Monetary policy and financial market developments in the US

Zivile Zekaite

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Adam Smith Business School
College of Social Sciences
University of Glasgow

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Abstract

Over the past decade, monetary policy has been in the spotlight as one of the key drivers of the real economy due to its aggressive response to the global financial crisis of 2007 - 2009. This has revived the debate of the late 1990s regarding the role of asset prices in policy decision making and has renewed interest in the impact of monetary policy on financial markets. Therefore, the focus of this thesis is the relationship between monetary policy conduct and financial market developments in the United States (US) over the period spanning the Great Moderation, the global financial crisis and its aftermath. Three empirical chapters analyse different aspects of monetary policy interaction with financial markets using alternative methodologies.

The first empirical chapter provides a comprehensive study of conventional monetary policy in the US. It investigates the Federal Reserve's response to financial market stress during the Great Moderation and the part of the global financial crisis by addressing two main questions. Firstly, does the Federal Reserve (Fed) react directly to the indicators of financial stress and, if so, is such reaction symmetric? Secondly, does the policy response to inflation and output gap change in light of financial turmoil? These questions are examined with respect to the four different dimensions of financial market stress: credit risk, stock market liquidity risk, stock market bear conditions and poor overall financial conditions. In addition, the analysis separately evaluates the impact of the latest crisis on US monetary policy. The results indicate the direct policy reaction to developments in the stock market price index, an interest rate spread, the measure of stock market liquidity and broad financial conditions that is found to be strongly dependent on the business cycle. Financial market developments have much more weight on the Fed's decisions during economic recessions as compared to economic expansions. Furthermore, in times of elevated financial distress, the Fed's reaction to inflation declines to some extent, while the output gap parameter becomes statistically insignificant. Nevertheless, the finding that financial stress implies a lower policy rate appears to be largely driven by monetary policy actions during the period 2007 - 2008. Thus, the financial crisis has had important implications for US monetary policy.

Chapter 2 investigates what explains the variation in unexpected excess returns on the 2-, 5- and 10-year Treasury bonds and how returns respond to conventional and unconventional monetary policy in the period spanning the Great Moderation, the recent

financial crisis and its aftermath. In addition, unexpected excess returns are decomposed into three components related to the revisions in rational market expectations (*news*) about future excess returns, inflation and real interest rates to identify the sources of the bond market response to monetary policy. The main findings imply that news about future inflation is the key factor in explaining the variability of unexpected excess Treasury bond returns across the maturities. Regarding the effect of conventional and unconventional monetary policy actions, monetary easing is generally associated with higher unexpected excess Treasury bond returns. Furthermore, the results highlight the importance of the inflation news component in explaining the reaction of the bond market to monetary policy. The positive effect of monetary easing on unexpected excess Treasury bond returns is largely explained by the corresponding negative effect on inflation expectations. Nevertheless, the bond market reaction to conventional policy shocks has grown weaker over the more recent period, perhaps reflecting changes in the implementation and communication of the Fed's policy since the middle 1990s. Meanwhile, the results with respect to unconventional monetary policy are driven to a great extent by the peak of the financial crisis in autumn of 2008.

Finally, Chapter 3 aims to revisit the role of conventional Fed's policy in explaining the size and value stock return anomalies, while taking fully into account the bi-directional relationship between monetary policy and real stock prices. As interest rate-based policy is of main interest here, the sample period ends prior to the crisis in 2007. The results confirm a strong, negative and significant monetary policy tightening effect on real stock prices at both aggregate and disaggregate (portfolio) levels. Furthermore, there is the evidence of the "delayed size effect" of monetary policy actions. Following a contractionary monetary policy shock, an immediate decline in stock prices of large firms is more pronounced as compared to small firms. However, large stocks recover to a great extent in the second period after the shock, while small stocks drop sharply. Meanwhile, the findings overall are not very supportive of the differential impact of monetary policy on value versus growth stocks as predicted by the credit channel. Finally, the results do not indicate the strong Fed's reaction to stock price developments.

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Author’s Declaration

I declare that, except where explicit reference is made to the contribution of others, that this dissertation is the result of my own work and has not been submitted for any other degree at the University of Glasgow or any other institution.

Signature:

Printed name: Zivile Zekaite

Introduction

In order to achieve the dual mandate of price stability and the maximum level of employment, the Federal Reserve (Fed) typically conducts monetary policy by setting the target level for the federal funds rate, i.e. an overnight interest rate at which depository institutions lend reserve balances held with the central bank to other depository institutions. Initially, policy rate changes influence other market interest rates. Generally, monetary policy tightening increases interest rates, although the effect at the long end of the yield curve is typically weaker (Evans and Marshall, 1998; Kuttner, 2001). According to the standard interest rate channel of transmission, contractionary monetary policy increases the real cost of borrowing and both consumption and investment spending decline. The link between market interest rates and asset prices, such as of stocks, bonds, and currency, enables monetary policy to have additional effects on aggregate demand through the wealth and credit channels (Mishkin, 2001; Kuttner and Mosser, 2002).

The dramatic and widespread impact of the global financial crisis of 2007 - 2009 on the functioning of financial markets and on the real economy prompted an aggressive response by the Fed as well as other major central banks. The federal funds target rate hit the zero lower bound, while various liquidity facilities were launched to reduce strains in financial markets. After exhausting conventional monetary policy tools, the Fed initiated the large-scale outright purchases of longer-term assets, mainly Treasuries and agency-guaranteed mortgage-backed securities, with an aim to reduce longer-term interest rates and to fulfil the dual mandate. Consequently, the federal funds rate has fallen short of being an adequate measure of monetary policy stance. Instead, quantity-based measures, such as the monetary base, bank reserve balances or central bank's assets, have been widely used to gauge unconventional policy actions. The transmission channel associated with these outright purchases, also known as quantitative easing, likely works through changes in relative asset prices due to central bank-induced changes in the outstanding quantities of these assets available to the public (Kuttner and Mosser, 2002).

This thesis examines the relationship between the Fed's policy decisions and financial market developments. It considers both the role of asset prices in setting the policy rate and the impact of monetary policy actions on the two key financial assets, i.e. government bonds and stocks. Motivated by the events in the period 2007 – 2008 and respective policy actions by the Fed, Chapter 1 is focused on the impact of financial market stress on setting the policy interest rate. It has been noted by some that the Fed may be responding to developments in financial markets in an asymmetric manner, i.e. easing policy stance in response to worsening financial conditions but being relatively unresponsive to upside developments in financial markets, such as the build-up of stock price bubbles (Neely, 2004; Roubini, 2006; Kahn, 2010). Others show that a different policy framework with respect to standard macroeconomic variables may be followed in times of intense financial distress (Alcidi, Flamini and Fracasso, 2011; Gnabo and Moccero, 2015). The motivation for the empirical analysis in Chapter 1 also stems from somewhat mixed empirical evidence of the Fed's reaction to financial market developments. A great number of studies find support for the Fed's response to asset price movements, mostly stocks, interest rate spreads and broader financial conditions (Chadha, Sarno and Valente, 2004; Alcidi, Flamini and Fracasso, 2011; Baxa, Horvath and Vasicek, 2013). Recent studies also indicate that the global financial crisis had a significant impact on the Fed's policy reaction function (Baxa et al., 2013; Belke and Klose, 2013). Meanwhile, others demonstrate that there is no significant reaction to asset price developments or broad financial conditions over and above their impact on expected inflation and output (Bernanke and Gertler, 1999; Fuhrer and Tootell, 2008; Castro, 2011).

Following the implementation of quantitative easing, the vast amount of empirical literature turned to evaluate its effects on longer-term interest rates as well as other financial assets. Typically, central bank's asset purchases are found to be effective in reducing longer-term interest rates. The existing literature identifies two key channels of transmission that explain the decline in long-term yields: the signalling channel (Christensen and Rudebusch, 2012; Bauer and Rudebusch, 2014) and the portfolio balance channel (Gagnon et al., 2011; D'Amico et al., 2012). Nevertheless, the empirical evidence as to which channel is more important is rather mixed indicating that the understanding of how quantitative easing led to lower bond yields is still incomplete. Consequently, Chapter 2 investigates the sources of variability in unexpected excess government bond returns and estimates the impact of both standard monetary policy and unconventional policies on bond returns and the components of these returns. As the majority of related studies employ either an event study, a structural VAR model or various term structure models, this chapter takes an alternative approach.

As the recovery of the real economy has eventually gained a strong momentum, the Fed has started to normalise its policy by raising the federal funds rate target for the first time in nearly a decade at the end of 2015. Thus, unconventional policies are to be gradually phased out and interest rate-based policy regains its importance. Given the prominent role of asset prices in the monetary policy transmission mechanism, especially that of stock prices, Chapter 3 examines the effect of conventional monetary policy on stock returns and revisits the role of the Fed's policy in explaining the size and value stock market anomalies. In the literature, it is generally found that monetary policy tightening depresses stock returns and the impact is stronger for small firms than for large ones and for firms with a high book-to-market ratio (value stocks) as compared to firms with a low book-to-market ratio (growth stocks). Nevertheless, the evidence for this differential stock returns response to monetary policy shocks as implied by the credit channel appears to be weaker and mixed since the 1980s (Guo, 2004; Tsai, 2011; Kontonikas and Kostakis, 2013; Maio, 2014). Also, previous studies are mostly focused only on one side of the potentially bi-directional relationship between monetary policy and stock prices. Meanwhile, the empirical studies that do account for this simultaneous relationship typically examine the aggregate stock market.

The first chapter begins with a literature review on the Taylor rule (Taylor, 1993) and outlines some important developments and main caveats encountered when estimating interest rate rules. The empirical analysis is based on the estimation of several alternative specifications of a forward-looking augmented Taylor rule using data for the period 1985:Q1 - 2008:Q4. Chapter 1 contributes to the literature by providing a comprehensive study of the Fed's monetary policy conduct with respect to financial market developments. The analysis considers four dimensions of financial distress: credit risk, stock market liquidity risk, stock market bear conditions and overall financial conditions. The impact of aggregate stock market liquidity conditions on the monetary policy interest rate has not yet been considered in the relevant literature. The chapter aims to answer two main questions. Firstly, does the Fed react to the indicators of financial stress and, if so, is such reaction symmetric? Secondly, does the policy response to inflation and output gap change in the presence of intense financial stress? Thus, the direct and indirect reaction of the Fed to financial variables is considered. This chapter also adds to the literature by providing an insight into how the past episodes of financial turmoil compare to the most recent financial crisis. The impact of the latest crisis on the main findings is examined relative to the past episodes of financial distress.

The results in Chapter 1 provide support for the direct policy reaction to developments in the stock market, the interest rate (credit) spread, stock market liquidity and broad financial conditions; however, it is found to be strongly dependent on the business cycle. Specifically, financial market developments have much more weight on the Fed's interest rate decisions in economic recessions as compared to the periods of economic expansions. On the other hand, this result is likely to be driven by the end of the sample period. With respect to the indirect policy response, during elevated financial stress, the Fed's reaction to expected inflation declines to some extent, while the output gap parameter becomes statistically insignificant. The indirect response to financial market stress becomes more evident in 2007 - 2008. The parameter on expected inflation declines significantly, turns negative and statistically insignificant. With respect to the output gap, the estimated coefficient increases slightly, but not substantially, and remains significant. Overall, the finding that financial stress implies a lower policy rate appears to be largely driven by the Fed's actions over the period 2007 - 2008. Thus, the latest crisis had a significant impact on the monetary policy framework with the focus shifting away from price stability towards the functioning of the financial system and financial stability.

Chapter 2 rests upon two strands of literature. The first one examines bond market determinants, such as macroeconomic factors, while the second one is focused on the effects of monetary policy actions, including quantitative easing, on the term structure of interest rates. This chapter adopts the log-linear approximation to the standard present value framework in a combination with a vector autoregressive (VAR) model (Campbell and Ammer, 1993) to investigate what explains the variation in unexpected excess returns on the 2-, 5- and 10-year US Treasury bonds over the period 1985:1 – 2014:2 using monthly data. Unexpected excess returns are decomposed into three components related to the revisions in rational market expectations (*news*) about future excess returns, inflation and real interest rates. In the spirit of Bernanke and Kuttner (2005), Chapter 2 identifies the sources of the bond market response to conventional and unconventional monetary policy. The contribution of the analysis in this chapter is three-fold. Firstly, the empirical approach allows explaining the bond market reaction to monetary policy changes in terms of news about macro-fundamentals, such as real interest rate and inflation, and expected excess bond returns, i.e. the risk (term) premium. This set-up has been largely overlooked in the bond market literature. Secondly, special attention is paid to the role of the financial crisis and unconventional policy subsequently adopted by the Fed. This is the first attempt to analyse quantitative easing effects within the VAR-based returns variance decomposition framework. Finally, shorter maturities are also considered, in addition to the commonly analysed 10-year Treasury bond.

The main findings of Chapter 2 show that news about future inflation is the key factor that drives the variability of unexpected excess Treasury bond returns across the different maturities. Regarding the effect of conventional and unconventional monetary policy actions, monetary easing is generally associated with higher unexpected excess Treasury bond returns, i.e. lower bond yields. Furthermore, the results highlight the importance of inflation news in explaining the bond market reaction to monetary policy. The positive effect of monetary policy easing on unexpected excess Treasury bond returns is largely explained by a corresponding negative effect on inflation expectations. Thus, the evidence is generally not supportive for the portfolio balance channel that implies a strong role for the risk (term) premium in explaining the response of bond yields to quantitative easing. Nevertheless, it is found that the reaction of bond returns to conventional policy shocks has become weaker over the more recent period, possibly reflecting changes in the implementation and communication of the Fed's policy since the middle 1990s. Meanwhile, the results with respect to unconventional monetary policy are driven to a great extent by the peak of the financial crisis in autumn of 2008.

Chapter 3 is focused on conventional monetary policy in the period of relatively favourable economic and financial conditions. To begin with, it reviews the empirical evidence from two typically employed methodologies to examine the policy impact on stock prices, i.e. event studies and structural VARs. The empirical analysis then investigates the effects of monetary policy on stock prices at aggregate and portfolio levels. Chapter 3 makes several contributions to the existing literature. Firstly, this chapter revisits the role of US monetary policy in explaining the size and value stock return anomalies within a structural VAR model based upon Bjornland and Leitemo (2009) that fully takes into account the simultaneous interaction between the policy rate and real stock returns. The model is identified using the combination of standard zero short-run restrictions and one long-run restriction that implies monetary policy neutrality. Hence, the contemporaneous relationship between real stock returns and the federal funds rate is unconstrained. Secondly, the original model specification as in Bjornland and Leitemo (2009) is augmented in line with the recommendations by Brissimis and Magginas (2006). Two forward-looking variables are included into the SVAR model: a market-based measure of expectations about the level of the monetary policy rate and a composite leading indicator of economic activity. This considerably improves the specification of the monetary policy reaction function and generates a sharper measure of monetary policy shocks. Finally, the main empirical analysis is conducted over the sample period, i.e. 1994:2 – 2007:7, that is not a standard choice in the SVAR literature. The motivation for the starting point stems from significant changes in the Fed's communication of policy decisions implemented at that time.

The results confirm a strong, negative and significant monetary policy tightening effect on real stock prices. Furthermore, there is the evidence of the "delayed size effect" of monetary policy actions. Following a contractionary monetary policy shock, the immediate decline in stock prices of large firms is more pronounced as compared to small firms. However, large stocks recover to a great extent in the second period after the shock, while at the same time small stocks drop sharply. The delayed response of smaller stocks to monetary policy shocks could possibly be explained as the result of their relative illiquidity and less frequent trading. Alternatively, the liquidity pull-back and portfolio rebalancing effects as well as the learning process of investors may play a role. With respect to the policy impact on value versus growth stocks, the value effect appears to be more evident only when firm’s size is controlled for. There is no evidence of stronger response of value firms as compared to growth firms to monetary policy shocks when ten value-sorted portfolios are considered. The evidence is more supportive of the credit channel when double-sorted size-value portfolios are considered instead. Within each size quintile portfolio, the most value stocks are more sensitive to changes in monetary policy conditions than the most growth stocks. Overall, the empirical findings provide some evidence, albeit not very strong, in favour of the credit channel of monetary policy transmission. Finally, the results do not indicate the strong policy reaction to stock price developments.

The remainder of this thesis is structured as follows. Chapter 1 reviews the literature on Taylor rules and provides the empirical study of the Fed’s reaction function. Chapter 2 discusses two strands of the literature relating to bond market determinants and monetary policy effects on the market interest rates. It then empirically examines the sources of the variability in Treasury bond returns and monetary policy impact on returns and their components. Finally, Chapter 3 reviews empirical evidence of conventional monetary policy impact on stock prices. The empirical analysis examines the simultaneous relationship between the Fed’s policy and real stock prices at market and portfolio levels.

Chapter 1: US monetary policy in times of financial market stress

1.1 Introduction

The global financial crisis of 2007 – 2009 posed serious challenges to monetary policymakers around the globe as it was of a greater order of magnitude as compared to the previous episodes of financial market distress. As nominal interest rates reached the zero lower bound in December 2008, the adoption of unconventional monetary policies, such as large-scale asset purchases by central banks, followed. The events during the crisis period have rekindled the academic debate of the late 1990s on whether the appropriate response of monetary policy to financial developments is proactive (Cecchetti et al. 2000) or reactive (Bernanke and Gertler, 1999; 2001). The pre-crisis consensus implied that monetary authorities should respond to asset price developments only to the extent they have implications for future inflation and output. It was argued against attempting to reduce or prevent asset price bubbles using monetary policy tools and many seemed to agree that mopping up after a bubble collapsed was a good policy (Greenspan, 2002; Blinder and Reis, 2005). However, this consensus appears to have shifted following the recent financial crisis and the argument goes that central banks should respond to financial imbalances independently of standard macroeconomic variables and the response should be symmetric (Wadhwani, 2008; Curdia and Woodford, 2010; Borio, 2014).

The global financial crisis was not the first time when the Federal Reserve conducted expansionary monetary policy in response to financial market distress. For instance, it eased monetary policy stance aggressively following the stock market crash in 1987 and 2000, the terrorist attacks in 2001, the Asian financial crisis in 1997 and the Russian default in 1998 (Neely, 2004; Roubini, 2006; Kahn, 2010). On the other hand, there is little evidence of a strong policy response to upside developments in financial markets, such as the stock price bubble in late the 1990s and the housing price bubble in the mid-2000s (Roubini, 2006). Thus, it appears that the Fed may be using a different monetary policy framework during the periods of high financial instability (Alcidi et al., 2011; Gnabo and Moccero, 2015). Nevertheless, the existing empirical evidence of the Fed's reaction to financial market developments is rather mixed. A vast number of studies find that the Fed sets its policy rate in response to asset price movements, mostly stocks, interest rate spreads and broader financial conditions (Chadha, Sarno and Valente, 2004; Alcidi, Flamini and Fracasso, 2011; Baxa, Horvath and Vasicek, 2013). Meanwhile, others demonstrate that there is no significant central bank's reaction to asset price developments over and above their impact on expected inflation and output (Bernanke and Gertler, 1999; Fuhrer and Tootell, 2008) and that the Fed does not consider broad financial conditions when deciding on the target rate (Castro, 2011).

Motivated by the above discussion, this chapter re-examines this conjecture. It begins with the discussion of the origins and development of the Taylor rule as well as some practical issues encountered in the literature on monetary policy rules. This chapter contributes to the existing literature by providing a comprehensive study of the Federal Reserve's response to financial market developments with respect to four different dimensions of financial market stress. In addition to the commonly considered types of financial distress, i.e. credit risk, stock market bear conditions and overall financial conditions, this chapter also examines the impact of aggregate stock market liquidity conditions on monetary policy decisions. To the best of my knowledge, this measure of financial market stress has not yet been considered in the related literature. The empirical analysis estimates several alternative specifications of an augmented forward-looking Taylor rule over the period 1985:Q1 - 2008:Q4. Two main questions are investigated. Firstly, does the Fed react directly to the indicators of financial stress and, if so, is such reaction symmetric? Secondly, does the policy response to inflation and output gap change in the presence of intense financial stress? Thus, a simple approach here considers both the direct and indirect reaction of the Fed to financial market developments and also tests whether this policy response is asymmetric. Furthermore, this chapter also adds to the literature by providing an insight into how the past episodes of financial turmoil compare to the most recent financial crisis. The empirical work to this respect for the US is relatively scant. The impact of the recent crisis is examined separately in an effort to evaluate how important it is for the overall findings.

The results provide support for both the direct and indirect monetary policy reaction to financial market developments. Nevertheless, this reaction appears to be largely driven by the Fed's actions in the period 2007 - 2008. While stock market returns, the credit spread, the measure of stock market liquidity and the financial conditions index are found to be statistically significant in the augmented Taylor rules, they only have a significant impact on the policy rate in recessionary periods. Moreover, it seems that the significant reaction to financial indicators during economic recessions can be explained, to a large extent, by the Fed's actions in response to the global financial crisis. With respect to standard macroeconomic variables, the Fed's reaction to expected inflation declines moderately in the periods of elevated financial stress, while the output gap parameter typically becomes statistically insignificant. Nevertheless, this indirect response to financial market stress strengthens considerably in the period 2007 - 2008. The parameter on expected inflation declines significantly, turns negative and statistically insignificant. With respect to output gap, the estimated coefficient increases slightly, but not substantially, and remains significant. Overall, the results imply a lower policy rate in times of severe financial market stress, especially during 2007 - 2008. As a result of the financial crisis, the Fed's policy framework appears to put much less emphasis on price stability and much more focus on financial market conditions. This view is in line with the evidence in Baxa, Horvath and Vasicek (2013) for the US and Martin and Milas (2013) for the UK.

The remainder of this chapter is set out as follows: Sections 1.2. and 1.3 provide the literature review. Section 1.4. outlines the methodology and the specifications of Taylor-type rules to be estimated. Section 1.5. describes the data. The discussion of the main empirical results is provided in Section 1.6. Robustness tests are discussed and summarised in Section 1.7., while Section 1.8. concludes.

1.2 The Taylor rule: origins and development

It is largely agreed that purely discretionary monetary policy is outperformed by rule-based policy in delivering better economic performance (Taylor, 1993; Orphanides, 2007; Taylor, 2012). Consequently, there have been numerous attempts in macroeconomic literature to describe the behaviour of monetary policy by a rule specified as an algebraic formula.[1] For instance, Milton Friedman's *k*-percent rule implies that a central bank should adopt the target of a steady money supply growth rate in order to achieve desired price stability on average (Friedman, 1968). Alternatively, McCallum (1988) suggests that monetary authorities should adjust the monetary base in response to nominal gross national product (GNP) deviations away from its target. Wicksell's simple interest rate rule, outlined in his work *Interest and Prices* in 1898, is briefly discussed in Orphanides (2007). It determines an interest rate as a policy instrument that is adjusted with respect to inflation, i.e. the interest rate is raised if price level rises and vice versa. However, it has no direct reference to real economic activity.

[1] Typically, such a policy rule is viewed more as a simple general framework used by monetary policymakers as a benchmark rather than as an ideal rule that is to be followed automatically.

As argued by Taylor (1993), good policy rules generally imply the reaction of policymakers to developments in inflation and real economic activity, i.e. a short-term interest rate is adjusted in response to changes in price level and real income. Taylor (1993) proposes the representative rule for the Fed that imbeds key principles outlined in the research of that time:

$$i_t = r^* + \pi_t + 0.5\left(\pi_t - \pi^*\right) + 0.5\hat{y}_t \tag{1.1}$$

where i_t is the nominal federal funds rate (FFR), r^* denotes the equilibrium real interest rate, π_t is the current rate of inflation over the previous four quarters, π^* is the target rate of inflation and $\hat{y}_t$ represents output gap, i.e. the percentage deviation of real gross domestic product (GDP) from its potential level at time *t*.[2]

According to Equation (1.1), the policy rate should be increased if either inflation exceeds its target level or real GDP is above its long-run trend. It also implies the "Taylor principle", i.e. the federal funds rate should be raised by more than an increase in inflation rate to guarantee an increase in real interest rate (Taylor, 1993). To see this, Equation (1.1) can be re-arranged as follows:

$$i_t = \alpha + 1.5\pi_t + 0.5\hat{y}_t \tag{1.2}$$

where $\alpha = r^* - 0.5\pi^*$.

Using vintage data on inflation and real GDP for the US, Taylor (1993) also demonstrates that this particular policy reaction function closely matches the actual FFR path in the period 1987 - 1992. Nevertheless, as Taylor (1993) notes, such a simple algebraic formula cannot and should not be followed mechanically, although it could serve as one of the inputs for monetary policy decision making and monetary authorities could use general principles underlying this policy rule. The Taylor rule has since become the cornerstone of the empirical work on monetary policy reaction functions.

On the other hand, it has been noted that outside the original sample period the standard Taylor rule fails to trace the actual funds rate as closely (Kozicki, 1999). Thus, the original specification has been augmented in two important dimensions to represent the

[2] Taylor (1993) sets the equilibrium real interest rate at 2% and the inflation target is also said to be equal to 2%.

behaviour of monetary authorities better.[3] Firstly, Equation (1.2) is based on a backward-looking measure of inflation and contemporaneous output gap. Given that there are time lags in monetary policy transmission, central bankers typically take pre-emptive actions towards developments in targeted economic variables implying a forward-looking policy reaction function (Clarida et al., 1998).[4] For instance, Orphanides and Wieland (2008) show that forward-looking rules based on economic forecasts by the Federal Open Market Committee (FOMC) explain the Fed's past decisions better than policy rules based on the observed data for the period 1988:Q1 – 2007:Q2. This finding is also consistent with the earlier study by Mehra and Minton (2007). They provide the evidence that the Fed is forward-looking with respect to inflation rate during the period 1987:Q1 – 2005:Q4.

The second dimension refers to interest rate smoothing. Equation (1.2) does not allow for gradual interest rate adjustments; however, it is typically a common practice among central banks to change their policy rate gradually.[5] Consequently, the standard Taylor rule has been augmented by including a lagged short-term interest rate as an independent variable (Evans, 1998; Judd and Rudebusch, 1998; Sack, 2000; Castelnuevo, 2003). For instance, Castelnuevo (2003) tests for interest rate smoothing in the forward-looking reaction function of the Fed over the period 1987:Q3 – 2002:Q3. The results are supportive of gradual federal funds rate adjustments. To address the argument by Rudebusch (2002), Castelnuevo (2003) also demonstrates that policy inertia does not appear to be induced by the misspecification of the estimated rule, i.e. the omitted variables problem. Generally, empirical studies allow for both forward-looking policy and interest rate smoothing when estimating monetary policy reaction functions (Clarida, Gali and Gertler, 2000; Mehra and Sawhney, 2010; Nikolsko-Rzhevskyy; 2011).

Taylor-type policy rules have attracted a great deal of attention in the literature and have been a good benchmark to examine past monetary policy decisions that may also provide a useful guidance for future policy making (Taylor, 1999; Kahn 2012). Nevertheless, one should have in mind several caveats when estimating interest rate policy

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[3] Note that there have been also some other adjustments, such as the inclusion of additional independent variables, that are discussed in more detail later in the text. With respect to two dimensions here, the literature has reached the consensus, while with respect to other developments the agreement is less evident.

[4] Batini and Haldane (1999) argue that forecast-based policy rules enable timely response to inflationary pressures leading to better control over price level and, possibly, also contribute to output stabilisation.

[5] There are several reasons why central banks may wish to smooth changes in the policy rate. Firstly, a gradual adjustment of interest rate reduces the possibility of causing disruption in financial markets due to potential overreaction to policy decisions. Secondly, policy decisions are better communicated to market participants allowing financial sector to anticipate future policy path. Thirdly, central banks might be uncertain about the exact state and structure of economy and, therefore, a gradual response is preferred in order to avoid policy reversals (Gerlach-Kristen, 2004). Furthermore, Sack (2000) suggests that inertial monetary policy could be explained by uncertainty surrounding monetary policy effects on economic variables.

rules. While many seem to agree that monetary policy is forward-looking and sets the policy rate in a smooth manner, there are other possible alterations to the standard Taylor rule. The following sections briefly discuss the main caveats that one should have in mind when estimating monetary policy rules.

1.2.1 *Ex post* versus real-time data

The interpretation of simple interest rate rules that are based on the *ex post* revised (final) data may provide with an inadequate policy description and recommendations. The informational issues arise due to the timeliness of the data available at the time of decision making. As macroeconomic data series are usually revised several times following their initial release, monetary policymakers do not have the *ex post* revised data, often used for estimations, at their disposal when deciding on policy actions. Orphanides (2001) compares the implied interest rates by the standard Taylor rule using both *ex post* and real-time series over the period 1987:Q1 – 1992:Q4. It is shown that the real-time data implies a lower interest rate as compared to the rate based on the revised data as used by Taylor (1993). Consequently, it is concluded that “the real-time policy recommendations differ considerably from those obtained with the *ex post* revised data” (Orphanides, 2001, p.965). In addition, the *ex post* revised data may obscure the fact that monetary policy is forward-looking. When real-time forecast data is used, Orphanides (2001) demonstrates that a forward-looking specification of the Fed’s policy rule has a better fit. In contrast, a contemporaneous policy rule specification is found to be the best using the *ex post* revised data.[6] In order to avoid these issues, more recent studies examining the Fed’s monetary policy typically employ real-time data (Boivin, 2006; Orphanides and Wieland, 2008; Mehra and Sawhney, 2010; Nikolsko-Rzhevskyy; 2011).

Nevertheless, some argue that the real-time data issue may not be as severe. For instance, Osterholm (2005) finds only minor differences between the estimated policy rules for the US using either the *ex post* or real-time data with respect to output during 1965:Q3 – 1999:Q4. Using both the real-time and final data for the US, Mehra and Minton (2007) obtain statistically significant estimates of response coefficients in the monetary policy rule for the period 1987:Q1 - 2005:Q4. The estimated coefficients are reasonably similar across the two types of data. Nevertheless, the estimated rules with the real-time data have a better fit. Furthermore, Molodtsova, Nikolsko-Rzhevskyy and Papell (2008) show that

[6] Alternative specifications with the respect to a horizon away from the decision period are estimated using either the *ex post* revised data with instrumental variables or, alternatively, the real-time forecast data with the ordinary least squares estimator.

there is no substantial difference between the estimated parameters for the US policy rule obtained with the revised and real-time data in the period 1979:Q1 – 1998:Q4.

1.2.2 Time-varying parameters

The ability of Taylor-type rules to adequately represent monetary policy decisions may also depend on the sample period. Larger implied policy rate deviations from the actual interest rate over certain periods indicate a potentially time-varying response of the Fed to inflation, output and, possibly, other factors (Judd and Rudebusch, 1998; Taylor 1999; Orphanides and Wieland, 2008; Kahn, 2010).

Taylor (1999) analyses the history of US monetary policy during the international gold standard era (1879 – 1914) and the Bretton Woods and Post-Bretton Woods period (1954 – 1997). The estimation of the standard Taylor rule confirms that the Fed's reaction to inflation and real economic activity has been changing over time. The interest rate response to inflation and output gap is found to be much weaker during the earlier sample period. In the Bretton Woods and Post-Bretton Woods eras, the estimated Taylor rule parameters continue to increase in magnitude over time and are more in line with the weights suggested by Taylor (1993) in the period 1987 – 1997 as compared to the estimates for 1960 – 1979. Furthermore, Taylor (1999) compares the rate implied by two basic policy rules, i.e. the standard Taylor rule and the Taylor rule with the output gap coefficient set to 1, with the actual FFR. The largest gaps between the actual and rule-prescribed rates are identified in the 1960s, 1970s and early 1980s. On the other hand, such deviations are relatively small in the period since the late 1980s and through the 1990s.

Similarly, Judd and Rudebusch (1998) examine the reaction function of the Fed over three periods associated with the three chairmen of the Fed: Arthur Burns (1970:Q1–1978:Q1), Paul Volcker (1979:Q3–1987:Q2), and Alan Greenspan (1987:Q3–1997:Q4). They demonstrate that the actual FFR is quite close to the policy rate prescribed by the standard Taylor rule in the Greenspan era. Meanwhile, the actual funds rate is constantly lower than the implied rate in the Burns period and is persistently higher than the recommended policy rate during the Volcker‘s chairmanship. The estimation of Taylor-type rules over the three periods reveals several important differences. The reaction function for the Greenspan era implies a strong and significant response of the Fed to both inflation and output deviations from their target levels, in line with the standard Taylor rule. Similar inferences can be made with respect to the Volcker period; however, the parameters are estimated with much less precision. In contrast, there is no evidence of the

Fed‘s reaction to inflation at the time of Burns‘s tenure; however, business cycle conditions appear to be quite important.[7]

Generally, the great part of the Greenspan-Bernanke era can be described by a relatively stable Taylor-type policy rule. Nevertheless, even over this period there have been misalignments between the actual monetary policy rate and the rate implied by alternative Taylor rule specifications. The rule-based era in 1985 – 2003 was followed by the period of significant deviations from the rule. Since the early 2000s until the onset of the global financial crisis, the federal funds rate was persistently lower than the prescribed policy rate, implying too accommodative monetary policy stance (Kahn, 2010; Taylor, 2012).[8] Moreover, Kahn (2010) identifies the funds rate's deviations from the rule-implied path also in the late 1990s. Following the financial crisis, the Taylor rule appears to prescribe a negative rate, while the actual rate is stuck at the zero lower bound (Kahn, 2012).

Several studies have attempted to explain such deviations from the rule-based path. For instance, Orphanides and Wieland (2008) demonstrate that using real-time FOMC economic projections to estimate policy rules significantly reduces the gaps between the actual and implied policy rate in the period 1988 – 2007. Consequently, the Fed's policy can be represented by a stable Taylor-type rule. Mehra and Sawhney (2010) analyse the period 1988 – 2006 and show that deviations from the policy rule largely disappear once changes in the policymakers' choice of inflation measure are accounted for. Alternatively, monetary policy deviations from the rule-based path may be explained by a central bank's reaction to factors not reflected in the policy rule, such as financial instability and asset prices, or by a decline in the equilibrium real interest rates (Hofmann and Bogdanova, 2012).

1.2.3 Measurement of independent variables

Another practical issue with estimating monetary policy rules is related to the measurement of variables included in a central bank's response function. With respect to macroeconomic data, a great variety of data series and methods of calculation may be used to obtain measures of, for instance, inflation and output gap. As empirical studies show, using alternative data series may result in a wide range of the implied policy rate by the estimated Taylor-type policy rules. For instance, Kozicki (1999) demonstrates that the

[7] Similar results are also reported in Seyfried and Bremmer (2001).

[8] Hofmann and Bogdanova (2012) discuss that this phenomena of the policy rate being below the rule-implied rate since the early 2000s is also true in the global context.

estimated parameters of policy rules for the US in 1983 – 1997 and 1987 – 1997 depend on the choice of inflation and output gap measures.[9] Depending on the measure of inflation, the equilibrium real interest rate may also vary over time and, in turn, may have implications for the policy rule recommendations (Kozicki, 1999). Similarly, many other studies compare the estimated policy rules using alternative series of macroeconomic variables, including unemployment gap measure or unemployment rate instead of output gap, and show that there are differences in the estimated parameters across specifications (Evans, 1998; Mehra and Minton, 2007; Orphanides and Wieland, 2008; Mehra and Sawhney, 2010; Kahn, 2012).

Nevertheless, if one allows for time-varying properties in the measures of relevant variables, the estimated policy rule could still describe the actual policy quite well. For instance, the Fed has changed the choice of the measure of inflation under consideration for policy making in February 2000. After this change is accounted for, it is possible to describe the historical US monetary policy with a stable forward-looking Taylor-type policy rule over the period 1987:Q1 – 2004:Q4 (Mehra and Sawhney, 2010).

1.2.4 Non-linearity in monetary policy rules

The standard Taylor rule is a linear reaction function, associated with the quadratic loss function of a central bank and a linear aggregate supply curve, and imposes equal weights on the devations of output and inflation away from their target levels independently of the size or direction of such deviations (Svensson, 1997). Nevertheless, monetary authorities may be responding to the target variables in a more complex way than suggested by a simple linear rule. For instance, Dolado, Maria-Dolores and Ruge-Murcia (2004) derive the optimal monetary policy rule using the model that departs from the standard framework in two important ways. Firstly, they allow for asymmetric central bank‘s preferences. Secondly, the aggregate supply relation is assumed to be convex, i.e. there is a non-linear relationship between inflation and output gap.[10] The resulting policy rule is the non-linear reaction function of a central bank that implies different weights on

[9] In this study the following measures of inflation are considered: CPI inflation, core CPI inflation, GDP price inflation, and expected inflation from the Survey of Professional Forecasters. With respect to the output gap, the alternative measures of potential real GDP are obtained from the Congressional Budget Office, the International Monetary Fund, the Organisation for Economic Cooperation and Development, and Standard and Poor’s DRI.

[10] Earlier studies extended the standard framework by relaxing the assumption about symmetric preferences of a central bank (Cukierman and Gerlach, 2003; Nobay and Peel, 2003). Alternatively, others assumed non-linear aggregate supply curve, i.e. inflation is a convex function of unemployment and, through the Okun’s law, the relationship between inflation and output gap is also non-linear (Schaling, 2004; Dolado, Maria-Dolores and Naveira, 2005).

the positive and negative deviations of inflation and output from their respective targets (sign asymmetry) and implies non-linear changes in interest rate response to changes in inflation and output gap (size asymmetry). Similarly, Cukierman and Muscatelli (2008) allow for changes in the nature of asymmetric central bank's preferences across different monetary policy regimes, i.e. recession-avoidance preferences prevail in normal times and inflation-avoidance preferences dominate in inflation-targeting period. Meanwhile, Florio (2009) builds the theoretical model that allows for asymmetries with respect to upward and downward movements in the policy rate.

Furthermore, a great number of empirical sudies provide the evidence of asymmetric monetary policy behaviour. Dolado, Maria-Dolores and Ruge-Murcia (2004) estimate a Taylor-type rule for the US over the period 1970:1 – 2000:12. The results show that the policy rule of the Fed appears to be linear with symmetric inflation preferences prior to 1979. In the period after 1983, the behaviour of the Fed is better described with a non-linear Taylor rule implying an asymmetric respone to positive and negative inflation gaps with the former having a greater weight. Ahmad (2016) investigates the Fed's policy over the period 1983:Q3 – 2007:Q4. The results show that the response to inflation is strong in the periods of output stability and inflation being outside the preferred range. Meanwhile, the Taylor principle is typically violated in the periods of distressed economic activity. Finally, in good economic conditions with price stability the Fed appears to be relatively unresponsive. Similar findings are also reported in Bunzel and Enders (2010).[11]

Bec, Salem and Collard (2002) show that in the post-1982 period the Fed fights inflationary pressures more aggresively during economic expansion, while recessionary periods are associated with output stabilisation. Similarly, Kazanas, Phillippoppoulos and Tzavalis (2011) also estimate a non-linear monetary policy rule for the US with respect to the business cycle. The results for the period 1960:Q1 – 2010:Q2 indicate that the Fed follows the Taylor rule in economic expansions with a positive and significant response to both inflation and output gap. However, the policy reaction to inflation falls sharply and turns insignificant during recessionary periods. Moeover, the response parameter to output gap also decreases and is only significant at 10% level.[12]

Furthermore, Florio (2009) provides the empirical evidence for asymmetric interest rate smoothing by the Fed in 1979:Q3 – 2005:Q4. During the Greenspan era, the Fed is

[11] On the other hand, Castro (2011) demonstrates that the Fed's policy is likely to be described well by a forward-looking linear Taylor-type rule over the period 1982:10 – 2007:12.
[12] Similar evidence is also found for other two countries under consideration, the United Kingdom and Japan.

found to be more cautious about policy rate increases versus cuts with a higher degree of policy inertia, while the opposite is true in the Volcker period. [13,14]

1.2.5 Zero lower bound

The severe consequences of the global financial crisis reminded of another limitation with respect to simple interest rate policy rules. The zero lower bound (ZLB) constrained policymakers and the estimated Taylor-type policy rules prescribed substantially negative rates since the early 2009 (Hakkio and Kahn, 2014). Given that a nominal short-term interest rate cannot go far below zero, unconventional monetary policy tools have been heavily used in order to further ease monetary policy stance. Thus, the federal funds rate has lost its ability to reflect monetary policy stance adequately. As Kahn (2012) notes, in the absence of unconventional policies, if one was to follow a simple interest rate rule when the ZLB is binding, it could have led to persistent undershooting of inflation and economic activity targets. Alternatively, one may either adjust the implied policy rate downwards, increase medium-term inflation target or act more aggresively before reaching the ZLB and less aggresively when moving away from it (Kahn, 2012).

The standard Taylor rule does not take into account the possibility of being at the ZLB for an extended period of time. Therefore, the specification has to be adjusted accordingly to account for the zero lower bound when it comes to estimating interest rate-based policy rules (Taylor and Williams, 2010). For instance, Belke and Klose (2013) use a real interest rate as opposed to a nominal rate as a dependent variable in the estimated reaction function. They exploit the relation that quantitative easing in the US has provided stimulus to real economy through changes in inflation expectations and real interet rates. Alternatively, other studies construct a shadow federal funds rate in order to gauge overal monetary policy stance before and after the crisis (Lombardi and Zhu, 2014; Wu and Xia, 2016). Most importantly, this shadow rate incorporates both conventional and unconventional monetary policies. For instance, Hakkio and Kahn (2014) evaluate the actual Fed's policy stance by comparing the shadow federal funds funds rate, based on the Wu-Xia model, with the policy rate prescribed by the estimated Taylor-type rules. They conclude that unconventional monetary policy was not sufficiently expansionary at the early stages of the ZLB.

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[13] Florio (2006) provides the empirical analysis of non-linearities in the Fed's policy rule for the period 1979:Q3 – 2004:Q3. The results show that the Fed tends to react to inflation more aggressively if inflation rate is increasing and it appears to adjust the rate more gradually when the policy stance is tightened.

[14] For more studies on non-linear Taylor-type policy rules see also Assenmacher-Wesche (2006), Kim and Nelson (2006), and Surico (2007), among others.

To summarise, there are several caveats that one should take into account when estimating Taylor-type policy rules. Nevertheless, a simple interest rate rule still appears to be a relatively good description of the historical Fed's monetary policy, especially so in the Greenspan period, if one allows for interest rate smoothing and accounts for the forward-looking behaviour of monetary policy.

1.3 Financial market implications for the Taylor rule

1.3.1 The importance of financial markets for monetary policy

As it has been discussed above, the actual federal funds rate has not always been in line with the implied rate by Taylor-type rules over the past few decades. The funds rate was substantially and persistently below the level recommended by the Taylor rule in the periods 1998 - 2000 and 2002 – 2006. Too accommodative monetary policy stance of the Fed may have contributed to the build-up of financial imbalances leading to the global financial crisis (Kahn, 2010). In addition to previously discussed caveats, such policy rate deviations could possibly be explained, at least partially, through the response of monetary authorities to financial market developments not taken into account by the standard Taylor rule (Hofmann and Bogdanova, 2012). Several adverse events in financial markets in the past coincided with more expansionary monetary policy stance than would be justified by the standard rule. This implies that the Fed may operate in a different monetary policy framework during the periods of financial turmoil and instability that pose great risks to economic growth and price stability.

Roubini (2006) argues that the Fed's response to developments in asset prices has been asymmetric, i.e. accommodating severe downside risks but not leaning against excessive upward movements in asset prices.[15] For instance, the Fed provided support to financial markets in light of severe disruptions to the financial system, such as the stock market crash in 1987, the Asian financial crisis in 1997, the Russian default in 1998 followed by the collapse of the Long-Term Capital Management hedge fund, the terrorist attacks in 2001 and its aftermath.[16] On the other hand, the Fed did not react aggressively

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[15] Hayford and Malliaris (2005) demonstrate empirically that there may be a negative relationship between stock market overvaluation and the federal funds rate in the period 1987 – 2000.

[16] When the stock market crashed on the 19th of October in 1987, broad stock price indices plunged sharply and financial markets suffered from the combination of elevated stock price volatility, low stock prices and low liquidity in the banking system (Hafer and Haslag, 1988). In response, the Fed provided liquidity by lending and lowering the funds rate target by around 80 basis points in the following months (Neely, 2004). Following the default of the Russian government in August 1998, interest rate spreads between risky and safe

neither to the rising stock price bubble in the late 1990s nor to the housing price bubble in the first decade of this century (Roubini, 2006). The most recent example of aggressively accommodative monetary policy in response to financial developments is associated with the global financial crisis that started in August 2007. In the early stages of the crisis, the Fed cut the funds rate target seven times reducing it by 3.25 percentage points in less than a year to alleviate adverse conditions in financial markets. In addition, various facilities were launched to improve liquidity conditions, such as the Term Auction Facility, the Primary Dealer Credit Facility, and the Term Securities Lending Facility (Cecchetti, 2009). Soon after the Lehman Brother's bankruptcy in September 2008, the federal funds target rate reached its zero lower bound and unconventional monetary policy tools were introduced to provide further stimulus for the US economy.[17]

The global financial crisis has had several implications to the pre-crisis thinking about the science of monetary policy. Firstly, it has been recognised that financial sector developments have a much stronger impact on economic activity than previously thought. Secondly, the linear-quadratic framework of optimal monetary policy appears to be inadequate in financial distress. Also, it has become clear that price and output stability do not guarantee financial stability (Mishkin, 2011). Even prior to the crisis, it has been argued that financial imbalances, such as excessive credit growth, unsustainable financial leverage and large deviations of asset prices from their fundamental values, may build up even in the periods of relatively stable and sustainable inflation (Borio and Lowe, 2002; 2004). These imbalances could lead to financial instability once favourable market conditions are unexpectedly reversed.[18] Consequently, a high degree of uncertainty over future financial asset prices and economic outlook may increase financing costs, tighten

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assets shot up as international investors ran for safety. In the US, falling long-term Treasury yields and equity prices and rising volatility in the stock market prompted the Fed to reduce its policy rate (Neely, 2004). In the event of the terrorist attacks on 11th September in 2001, the loss of people and severe physical damage of the buildings of financial institutions forced the affected markets to close. The payment systems were significantly disrupted and financial market liquidity fell sharply. In addition to liquidity injections via open market operations and discount window lending, the Fed cut the policy rate target to 1.75% by the end of 2001down from 3.5% prior to September 11th. These later cuts in the policy rate are viewed as longer-term help for weak US economy (Lacker, 2004; Neely, 2004).

[17] For the more detailed discussion of these policies refer to Fawley and Neely (2013) and the literature review of Chapter 2.

[18] Firstly, higher asset prices directly influence the level of wealth and have an impact on consumption and investment spending through capital gains. Secondly, changes in asset prices are reflected in the balance sheets of firms and households, thus, increasing prices improve the creditworthiness of both and increase borrowing and spending. More wealth, consumption and better creditworthiness do not have adverse effects on real economy. Nevertheless, it may be detrimental to the economy when asset price bubble, unjustified by fundamentals, eventually bursts (Ferguson, 2005). Sharp declines in stock and housing prices, among other assets, reduce consumption and worsen balance sheet conditions since the value of collateral and capital declines substantially. Furthermore, lower asset prices increase a counterparty risk potentially leading to the systemic risk in the financial system. In turn, severe disruptions in financial intermediation may result into the second-round effects on consumption and investment (Neely, 2004).

credit conditions, and induce a sharp and long-lasting contraction in real economic activity and overall price level instability (Borio and Lowe, 2004; English, Tsatsaronis and Zoli, 2005; Hakkio and Keeton, 2009).

Due to the close nexus between financial and economic stability, financial market developments could be good indicators of future financial and economic conditions that policymakers may wish to consider in more detail. For instance, Borio and Lowe (2002) examine financial conditions in thirty four countries over three decades and find that substantial deviations in credit growth, stock and real estate prices, and investment from their respective trends are associated with future financial crises. They show that rapid credit growth together with large increases in asset prices increases the probability of financial instability and a crisis. Similarly, Schularick and Taylor (2012) evaluate the historical paths of money and credit growth in fourteen countries during 1870 - 2008. The results imply that strong and unsustainable credit growth in the preceding five years increases the likelihood of a financial crisis.[19]

Overall, there seems to be a strong rationale for considering financial market developments in a monetary policy reaction function (Roubini, 2006). The events in 2007-2009 have re-opened the debate with respect to the appropriate monetary policy response to asset prices and financial conditions in general. The dominant view prior to the crisis appeared to favour inflation targeting without a central bank's response to asset prices, except insofar they signal shifts in inflation expectations, as the best policy framework (Bernanke and Gertler, 1999; 2001). Nevertheless, extraordinarily high costs associated with the recent housing bubble bust in the US and the global crisis that followed may have shifted the consensus towards the opposite view that monetary policy should respond to asset prices and forming asset price bubbles, and it should do so in a symmetric manner (Wadhwani, 2008; Borio, 2014). The next section provides a more detailed discussion of this debate.

1.3.2 Should central banks react to financial market developments?

[19] It is a difficult task for monetary policymakers to monitor many aspects of financial market developments. To this respect, a composite indicator of financial stress and overall financial conditions could be very useful (English, Tsatsaronis and Zoli, 2005; Hakkio and Keeton, 2009; Carlson, Lewis and Nelson, 2012). For instance, sharp deviations of an index from its average level could imply that some preventive policy action is needed. Similarly, Brave and Butters (2012) show that the National Financial Conditions Index calculated by the Federal Reserve bank of Chicago appears to be a very good indicator and a powerful predictor of financial distress in the year ahead. Kliesen, Owyang and Vermann (2012) provide the survey of various indices of financial conditions and financial stress, i.e. FCIs and FSIs.

In one of the first papers on monetary policy response to asset prices, Bernanke and Gertler (1999) argue that central banks should disregard asset price movements and instead focus on inflationary pressures in the economy as such inflation-targeting framework helps to achieve both price and financial stability. They investigate what type of forward-looking monetary policy rules perform best in the presence of exogenous non-fundamental movements in asset prices.[20] The “inflation accommodating” policy assigns the reaction coefficient to expected inflation of just above one, while the “aggressive inflation targeting” approach sets the coefficient equal to two. Next, both policy rules are augmented to allow for a response to stock market developments. Model-based simulations indicate that aggressive inflation targeting without the reaction to stock price deviations from their steady state performs best in terms of minimising adverse effects of an asset price bubble on real economy (Bernanke and Gertler, 1999). Thus, monetary policy should consider asset prices only to the extent that they have implications for inflation expectations. Similarly, Bernanke and Gertler (2001) use the previously employed macroeconomic model to consider the entire probability distribution of shocks as opposed to only taking into account the worst outcomes. In line with their earlier work, the policy rule that implies the strong reaction to expected inflation and allows for the reaction to output gap but not asset prices, achieves the best outcome in reducing economic instability following asset price and technology shocks. Similar views are also expressed by Vickers (2000), Bullard and Schaling (2002) and Bean (2003).

The case against the monetary policy response to asset prices is also supported by arguments about the effects, identification and measurement of asset price bubbles. Firstly, it is argued that not every asset price bubble necessarily leads to the disruption of the financial system with adverse consequences for real economy (Bernanke and Gertler, 1999; Mishkin, 2008). Secondly, it is almost impossible to clearly distinguish between changes in asset prices due to fundamental factors and due to non-fundamental factors (Cogley, 1999; Vickers, 2000). Mishkin (2008) suggests that macroeconomic consequences of an asset price bubble may be limited provided that monetary policy becomes more accommodative following the bust of a bubble.[21] In the absence of certainty that an asset price bubble exists, monetary policy actions to deflate it may prove

[20] The simulations are based on the standard dynamic new-Keynesian model with financial accelerator effects that also allows for asset price bubbles.

[21] Generally, this view is in line with no policy reaction to a build-up of financial imbalances, while mopping-up once the bubble bursts. Nevertheless, others argue against such an asymmetric treatment of asset price developments (Roubini, 2006; Wadhwani, 2008). Furthermore, Borio (2014) notes that, in order to ensure monetary and financial stability, there should be a systematic and symmetric response of monetary policy to financial booms and busts. This implies leaning more against the financial imbalances in a boom phase and easing less aggressively and persistently when there is a bust.

destabilising and counterproductive (Cogley, 1999). Furthermore, the effects of interest rate changes on asset prices are not clear-cut (Mishkin, 2008).

Using the standard dynamic stochastic general equilibrium model (DSGE) augmented to allow for excess volatility in asset prices and macroeconomic variables, Gelain, Lansing and Mendicino (2013) examine the consequences of monetary policy response to financial imbalances. They find that the inclusion of house price inflation and credit growth directly into an interest rate rule tends to magnify inflation volatility. Furthermore, Assenmacher-Wesche and Gerlach (2010) estimate a forecasting model for inflation and output for eighteen countries over the period 1986 – 2008. The results show that the deviations of credit growth and asset prices away from their respective trends to contain little additional information with respect to future economic conditions once current inflation, output gap and interest rate are taken into account.

Contrary to the consensus of that time, Cecchetti et al. (2000) demonstrate that it is optimal for monetary policymakers to directly (and symmetrically) respond to asset prices in general as well as to asset price movements away from their fundamental values. Such policy response has become known as “leaning against the wind”. They augment the macroeconomic model used for the simulations by Bernanke and Gertler (1999) in several ways, such as computing optimal policy rules and allowing for interest rate smoothing. The simulation results clearly indicate that central bankers should act pre-emptively and react to asset price changes to deliver better macroeconomic performance and possibly reducing the likelihood of asset price bubbles. Furthermore, Cecchetti et al. (2000) argue that even though it is rather difficult to identify and measure asset price misalignments, the same is true for estimating some macroeconomic indicators, for instance, output gap. Thus, central banks should attempt to assess and respond to asset price misalignments. Similarly, Filardo (2004) concludes that the optimal monetary policy rule does include asset prices, although the policy response to non-fundamental movements in asset prices is preferred. Using a macroeconomic model allowing for wealth effects and financial market inefficiency, Kontonikas and Montagnoli (2006) also provide support for the systematic monetary authorities’ reaction to the deviations of asset values from their fundamental values as it minimises output and inflation volatility. Kontonikas and Ioannidis (2005) examine the interactions between asset prices and monetary policy within a structural rational expectations open economy model. They conclude that the response of a central bank to asset price misalignments is associated with lower macroeconomic variability.

The rationale for monetary policy to lean against financial imbalances has strengthened in recent years (Mishkin, 2011). Various theoretical studies show that it may be optimal to include the measures of financial conditions, such as interest rate spreads or credit growth, into Taylor-type policy rules as it increases welfare and macroeconomic stability by reducing a negative impact of financial instability on real economy. As Curdia and Woodford (2010) argue, aggressive policy rate cuts by the Fed between late 2007 and early 2008 are likely to reflect the policy response to financial sector stress. Using a DSGE model with credit frictions, they show that the inclusion of credit spreads, i.e. differences in borrowing interest rates across the classes of borrowers, into an otherwise standard Taylor rule could improve macroeconomic performance. This implies setting a lower policy rate in response to higher credit spreads. In addition, Gilchrist and Zakrajsek (2011) also demonstrate that allowing for the direct central bank's response to financial conditions, measured as changes in credit spreads, dampens a negative impact of financial instability on real economy.[22] Furthermore, Davis and Huang (2013) show that it may be optimal for policymakers to respond to movements in an interbank lending spread caused by an external financial sector shock, such as an increased financial risk or uncertainty. Alternatively, based on a theoretical macroeconomic model, Christiano et al. (2010) suggest that the modification of the Taylor rule by including a credit growth rate reduces the size of boom-bust cycles. In this case, monetary policy stance should be tightened when credit growth is very strong and loosened otherwise. According to Freixas, Martin and Skeie (2011), central banks should respond directly to financial distress in the banking sector. They develop a model of interbank market and argue that the policy rate should be lowered in response to distributional liquidity shocks to encourage the reallocation of liquid assets in the interbank market.[23]

To summarise, recent theoretical studies provide increasing support that monetary policy should respond to financial market developments. There is little doubt that severe disruptions to the functioning of financial markets often lead to dramatic consequences for real economy. Nevertheless, it is also typically agreed that the policy reaction to financial imbalances should be cautious and moderate given uncertainties with respect to asset price bubbles and the impact of monetary policy on asset prices and real economy.

1.3.3 Do central banks react to financial market developments?

[22] See also Taylor (2008) and Teranishi (2012), amongst others.

[23] For other recent theoretical studies that argue in favour of the central bank's response to asset prices see Pavasuthipaisit (2010), Gambacorta and Signoretti (2014), among others. On the other hand, Svensson (2016) argues that the costs of leaning against the wind may exceed the benefits.

As it has been discussed, there is a strong theoretical motivation for central banks to take into account financial market developments. Focusing on the US, this and the subsequent sections review empirical evidence from the estimated interest rate rules that allow for a policy response to financial variables. Nevertheless, the estimation of augmented Taylor-type rules is not the only method to gauge central bank's reaction to asset prices.[24]

Some initial empirical evidence is provided by Bernanke and Gertler (1999). Using the Generalised Methods of Moments (GMM), they estimate a forward-looking Taylor rule augmented with stock market returns for the Fed over the period 1979:10 - 1997:12.[25] Since stock returns are also included as instrumental variables, the monetary policy response to asset prices to the extent that they help predict future inflation and output gap is accounted for. The results show that the federal funds rate's response to stock returns is small, statistically insignificant and negative. These findings are also in line with the model-based simulations in Bernanke and Gertler (1999), indicating that the Fed does not react to stock prices over and above stock price implications for inflation and output forecasts.

Similarly, Chadha, Sarno and Valente (2004) examine the reaction of monetary policy to asset prices using dividend-price ratios for aggregate stock price indices and exchange rates for the United States, the United Kingdom, and Japan in the period 1979:9 – 2000:12. The GMM estimations of augmented forward-looking Taylor rules indicate several points. Firstly, the Fed is found to be responding to asset prices and the estimated parameter is positive and statistically significant. Secondly, the reaction coefficient to exchange rate, defined as a domestic price of foreign currency, is also positive and significant. This implies that the federal funds rate increases as stock prices increase and the exchange rate depreciates. This evidence suggests that asset prices are not treated solely as informational variables to forecast inflation and output developments. Similarly,

[24] For instance, Rigobon and Sack (2003) measure the Fed's reaction to stock market movements using a daily structural vector autoregression (VAR) model that is identified based on the heteroscedasticity of stock market returns. The identification of stock price shocks rests upon the observed shifts in the covariance matrix of reduced-form residuals and the assumption of homoscedastic monetary policy shocks. The results for the period 1985 – 1999 confirm a statistically significant Fed's reaction to stock market developments. Nevertheless, it has been suggested that the response is just enough to offset the impact of stock price shocks on economy. Using the same approach, Furlanetto (2011) also finds a positive and significant response of the Fed to the stock market. However, the response is found to be much weaker in the period 1988 - 1999. It declines further and becomes insignificant over the period 2003 - 2007. Others use lower-frequency VAR models identified with short- and long-run restrictions and find a significant positive response of the Fed to increasing stock and house prices in the post-1983 period (Bjornland and Leitemo, 2009; Bjornland and Jacobsen, 2013). Finally, Finocchiaro and von Heideken (2013) analyse the reaction functions of the Fed, BoE and BoJ within the estimated DSGE model. They find evidence for a significant and positive central banks' response to house price developments in 1983 - 2008.

[25] This estimation method allows using the actual (future) values of independent variables instead of their expected values. Instrumental variables known at time t-1 or earlier are chosen to empirically project theoretical expected values. This approach provides the consistent estimates of policy rule parameters.

the results are supportive of the direct, positive and significant response of the Bank of England (BoE) to stock prices and exchange rate movements. With respect to Japan, it is only the exchange rate variable that appears to matter.[26]

On the contrary, Fuhrer and Tootell (2008) argue that the Fed sets its policy rate with respect to stock market developments only insofar equity prices help to forecast policy goal variables. For the period 1966:Q1 – 2006:Q1, they estimate alternative interest rate rules augmented with stock market returns and compare the results across two different estimation approaches. Firstly, the GMM technique is used with the *ex post* data. The results indicate that (lagged) stock prices may play an independent role in monetary policy decisions in addition to their informational content with respect to future economic conditions. Nevertheless, this finding does not hold if information available to the Fed at the time of decision making is fully taken into account. The second approach involves estimating augmented Taylor-type rules using lagged stock prices and the Greenbook forecasts for inflation, real GDP growth and unemployment rate as instrumental variables. The estimated reaction coefficient to equity returns becomes insignificant.

Jovanovic and Zimmermann (2010) estimate augmented forward-looking Taylor rules for the US that include a proxy for stock market uncertainty based on the realised volatility, expected volatility, and implied volatility (VIX volatility index) over the periods 1991:1 – 2008:5 and 1980:9 – 1990:12. Generally, the GMM estimates indicate that there is a systematic response of the Fed to the level of uncertainty in the stock market. The interest rate response coefficient with respect to stock market volatility is typically negative and statistically significant implying a lower policy rate in times of greater financial instability. Moreover, according to the estimated augmented rules, the response of the Fed to expected inflation rate declines considerably as compared to the estimated rules without the stock market variable.[27]

Gerlach-Kristen (2004) finds that in the presence of financial stress the Fed tends to cut the interest rate more than what is warranted by inflation and output gap developments. Initially, the study examines whether the observed interest rate smoothing in the policy

[26] Using the GMM approach, Botzen and Marey (2010) find similar evidence for the European Central Bank (ECB). The results show a significant reaction of the ECB to stock price developments independently of their impact on future inflation and output during 1999 - 2005.

[27] In the related study, Bleich, Fendel and Rulke (2013) use expected implicit stock market volatilities to proxy for financial market stress and estimate augmented forward-looking Taylor rules for the central banks in the US, UK, euro area and Japan allowing for the direct policy response to financial distress. The findings imply a systematic (negative) response to expected stock market volatility by all monetary authorities except for the Bank of Japan. Thus, higher financial stress level triggers lower monetary policy rates in the US, UK and the euro area.

reaction function for the Fed can be explained by policy inertia or omitted variables.[28] Using the unobservable components technique, it is demonstrated that both the lagged interest rate and unobserved variable are statistically significant in the estimated monetary policy rule for the period 1987:Q4 – 2003:Q3. Furthermore, Gerlach-Kristen (2004) argues that this omitted unobservable variable could be related to financial market conditions. The estimation of an augmented Taylor-type rule reveals that the interest rate spread between the yields on the 10-year Treasury and Moody's BAA-rated corporate bonds enters the policy reaction function positively and significantly. This implies lower interest rates in response to more negative credit spreads, i.e. increasing financial market stress.[29]

Using the OLS, Borio and Lowe (2004) estimate alternative contemporaneous Taylor-type rules augmented with the indicators of financial imbalances for a group of countries over the period 1983:Q1 – 2002:Q2. To measure imbalances, they compute credit and equity gaps defined as the deviations of the ratio of private credit to GDP and real stock prices from their trends, respectively. In addition, they use a dummy variable to denote periods of potential banking distress when both gaps exceed their critical threshold. Overall, the results do not provide support for a significant monetary policy response to financial imbalances in Australia, Germany, and Japan. Nevertheless, the evidence of the monetary policy reaction to financial imbalances is somewhat stronger for the US, but only in the case without an interest rate smoothing term in the policy rule.

Alternatively to individual financial variables, several empirical studies employ the composite measures of financial conditions. For instance, Montagnoli and Napolitano (2005) examine the forward-looking policy reaction functions for Canada, the euro area, the UK and the US. Firstly, a financial conditions index (FCI) is constructed for each region with the focus on three asset prices: real exchange rate, real house prices and real stock prices. Secondly, the FCI is then included in a respective forward-looking Taylor-type rule for each country as an additional indicator of future developments in economic and financial market conditions. The GMM estimations indicate that the volatility of residuals of the augmented Taylor rules is typically smaller as compared to the standard Taylor rule. Furthermore, for the period 1985:5 – 2005:5, the FCI enters the estimated rule positively and significantly in the US, the UK and Canada. This implies that central banks

[28] The interest rate smoothing term may not only represent a gradual adjustment of the short-term rate but also some serially correlated variables that are incorrectly excluded from the estimated regression (Rudebusch, 2002).

[29] Similarly, Alcidi, Flamini and Fracasso (2011) also provide the empirical evidence of the Fed's reaction to the spread between the Moody's BAA corporate bond index and 10-year Treasury bonds in the period 1987:Q3 – 2005:Q4. If credit spread widens, the policy rate declines.

may be responding to developments in financial conditions in addition to their reaction to standard policy target variables.

Similarly, Castro (2011) constructs FCIs in the spirit of Montagnoli and Napolitano (2005) and estimates the augmented forward-looking Taylor rules for the ECB, the Fed and the BoE. The GMM approach is used to estimate the rules augmented with the financial conditions index for each country. With respect to the US, the results for the period 1982:10 – 2007:12 show that there is no significant reaction to financial conditions by the Fed. On the other hand, the credit spread between long-term risk-free government bond and corporate bond yields (as deviations from trend) enters the Taylor rule with a positive and significant parameter. In contrast to the results reported by Montagnoli and Napolitano (2005), the ECB is found to be responding to financial conditions, while this is not the case for the BoE.

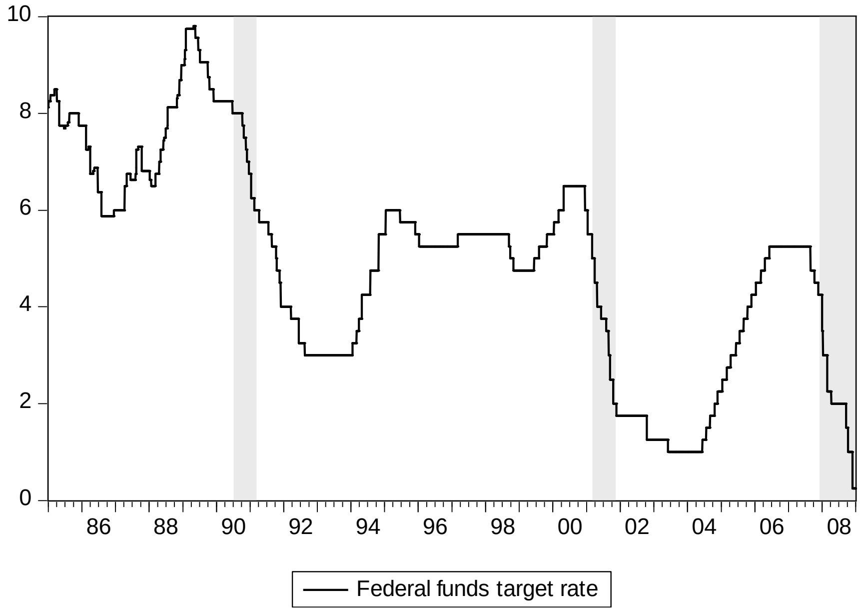
1.3.4 Do central banks respond differently in bad financial conditions?

The evidence discussed above is based on the estimation of linear policy rules and is somewhat mixed. Nevertheless, it may be argued that the response of monetary authorities to financial imbalances is likely to be asymmetric. In times of financial distress, monetary policymakers appear to put more emphasis on financial market developments in order to stabilise real economy. In contrast, their reaction to the build-up of financial imbalances, for instance, stock price bubbles, is not as evident (Roubini, 2006; Mattesini and Becchetti, 2009). One of the reasons for such an asymmetric behaviour may be the interaction between financial conditions and real economic activity. For instance, Hubrich and Tetlow (2015) find that the relationship between macro economy and financial stress is non-linear, i.e. shocks to the financial system have more damaging effects on output in times of financial stress as compared to normal times. Some anecdotal evidence is presented in Figure 1.1. Typically, federal funds target rate cuts are somewhat more aggressive than rate increases. As monetary policy expansion is usually associated with economic recessions and financial instability, Figure 1.1 suggests that the Fed's policy may be asymmetric with respect to economic and financial conditions.

Recent literature on monetary policy rules provides empirical evidence in favour of a different monetary policy framework in the periods of intense financial stress.[30]

[30] Note that it may not only be the response of the policy rate to financial variables that is non-linear, but it also may be the response to standard goal variables that could vary with the level of financial stress.

Figure 1.1: Federal funds target rate



Notes: This figure plots daily time series for the federal funds target rate over the period 1985 – 2008. The shaded areas denote the periods of US recessions as defined by the NBER.

For instance, Mattesini and Becchetti (2009) show that the Fed tends to cushion declining stock prices below their fundamental values with policy rate cuts, while it does not respond to the overvaluation of stocks. They develop the measure of stock price misalignments, i.e. the Index of Stock Price Misalignment (ISPM), and include it into the Fed's reaction function. The augmented forward-looking Taylor rules are estimated using the GMM for the period 1980:1 – 2001:4. They find mixed evidence with respect to the symmetric Fed's response to the ISPM. Next, they distinguish between the positive and negative values of the measure of stock price misalignments. The results show that the estimated parameter for the negative values of the ISPM is positive and statistically significant. In contrast, there is little evidence of a significant response by the Fed to the positive values of the ISPM (Mattesini and Becchetti, 2009). This implies that the Fed responds to the undervaluation in the stock market by conducting expansionary monetary policy, but it largely ignores potential stock price bubbles.[31]

Similarly, Hoffmann (2013) investigates whether the Fed and ECB respond asymmetrically to stock price inflation deviations from the trend. The GMM estimations indicate that the Fed cuts the policy rate target in response to declining stock market, but it does not increase the target to dampen rising stock prices in the Greenspan-Bernanke

[31] In addition, Ravn (2012) also provides the empirical evidence that the Fed responds asymmetrically to asset prices during the period 1998 - 2008. Using the methodology of Rigobon and Sack (2003), it is demonstrated that the Fed ignores increasing stock prices, but it reduces its policy rate target in response to declining stock market.

period, i.e. 1987:8 – 2008:12. On the other hand, the ECB is not found to react to stock market developments, but it does respond to developments in the exchange rate over the period 1999:1 – 2008:12. The ECB decreases the policy rate in response to the appreciation of the euro with respect to the US dollar and vice versa.

With respect to the US and Japan, Borio and Lowe (2004) present some empirical evidence of an asymmetric response to credit growth and asset price deviations from their respective trends in the period 1983:Q1 – 2002:Q2. The results of the estimations of augmented Taylor rules indicate that in response to adverse financial imbalances, i.e. negative credit and equity gap, policymakers tend to relax monetary policy stance by more than what is suggested by inflation and output gap tendencies. Meanwhile, there is typically no significant reaction to positive credit and equity gaps. This translates into lower policy rates when financial conditions worsen. Nevertheless, this finding only seems to be valid in the Taylor rule specifications without an interest rate smoothing term.

Belke and Klose (2010) demonstrate that the monetary policy framework of the Fed changed considerably around the crisis in 2007 - 2009. They estimate alternative Taylor rules augmented with financial variables prior to the crisis (1999:1 – 2007:1) and during the crisis (2007:8 - 2009:6) with the GMM. Across the specifications, the Fed's response to inflation is positive and statistically significant in the pre-crisis period, albeit the coefficient is smaller than one. In contrast, the reaction to inflation becomes weaker during the crisis and the estimated parameter is mostly negative and significant. With respect to output gap, the estimated parameter decreases somewhat in magnitude during the crisis, but it is typically positive and statistically significant in both sample periods. Prior to the crisis, the Fed is found to respond to commercial and industrial credit and money growth with the estimated coefficients being positive and statistically significant. This implies a lower policy rate when credit and money growth declines. The Fed also appears to decrease the federal funds rate target if the interest rate spread between long-term and short-term government securities is higher, indicating a rising risk in capital markets. Interestingly, the response parameters to stock and house price inflation are both negative and statistically significant. In the crisis period, the estimated parameters with respect to additional variables change the sign, except for the interest rate spread, and remain statistically significant. For instance, the response to asset price inflation turns positive. Thus, the policy rate is reduced if asset prices fall during the crisis; however, it is reduced in response to rising stock and housing prices prior to the crisis.

Similarly, Belke and Klose (2010) also analyse the reaction function of the ECB. Before the crisis hit, the ECB's reaction to inflation and output gap is denoted by positive and significant parameters. During the crisis, the response to inflation appears to have increased, while the reaction to output gap has decreased and even turned negative. Furthermore, the ECB's reaction to financial variables is typically statistically significant. Nevertheless, some coefficients differ in size and sign as compared to their respective pre-crisis sample estimates. For instance, there is some evidence of more aggressive policy easing in response to declining credit growth and increasing interest rate spread in the crisis period than prior to the crisis. Moreover, there is some indication of the asymmetric ECB's response to asset price inflation. In the crisis period, the ECB seems to have reduced the policy rate when housing prices rose, while the opposite response to increasing housing prices prevailed before the crisis.

Belke and Klose (2013) extend their earlier work by allowing for monetary policy inertia and by taking into account the ZLB period, i.e. a real short-term interest rate is used instead of a nominal rate as a dependent variable. With respect to both the Fed and the ECB, they find that the interest smoothing parameter has decreased during the financial crisis implying more aggressive interest rate setting policy. For the Fed, the results are broadly in line with those reported in Belke and Klose (2010). Firstly, the reaction to inflation falls sharply during the crisis and even turns negative for some specifications. The estimated parameter of monetary growth turns from positive and significant before the crisis to negative and significant afterwards, indicating that inflationary pressures become less important during the financial crisis. The Fed is found to respond to credit growth only after the onset of the crisis, although the estimated coefficient is negative and significant. The response to an interest rate spread is somewhat smaller during the crisis, but it is always negative and significant. Finally, in the crisis period the Fed appears to have lowered the policy rate in response to falling asset prices while accommodating asset price booms before the crisis. With regards to the ECB, the results are somewhat different as compared to Belke and Klose (2010). The inflation parameter turns negative in the majority of specifications in line with the findings for the Fed. Similarly, the ECB reacted more aggressively to asset price inflation and somewhat less so to interest rate spreads in the crisis period than before the crisis. In contrast to the Fed, the credit growth and money growth parameters are found to be negative and significant prior to the financial crisis, but they turn positive and significant during the crisis.

The monetary policy reaction function for the UK also appears to have changed during the financial crisis. Martin and Milas (2013) examine alternative Taylor-type rules for the period 1992:10 – 2010:7. The findings from the GMM estimations show that prior to the financial crisis the BoE's policy can be described by a simple forward-looking

Taylor rule. On the contrary, the response to inflation almost disappears in the crisis period as the parameter becomes negative and insignificant in 2007:5 – 2010:7. Meanwhile, the response parameter to output gap also decreases, but it remains positive and significant. In the next step, they estimate policy rules augmented with either the IMF financial stress index for the UK or the Federal Reserve Bank of Kansas City Financial Stress Index. Both indices are insignificant prior to the crisis; however, they appear to play a dominant role in interest rate setting decisions in the crisis period. Higher financial stress is associated with a lower monetary policy rate. Furthermore, Martin and Milas (2013) develop a smooth-transition model that describes full-sample policy rate decisions more adequately. They show that the non-linear monetary policy rule, determined as a weighted average of crisis and no-crisis policy regimes, best explains the UK monetary policy.[32] The model implies that the BoE follows the conventional Taylor rule outside the crisis regime, while the interest rate is mainly set in response to the financial stress index and output gap in the crisis regime.

Similarly, Baxa, Horvath and Vasicek (2013) investigate the role of financial stress for interest rate setting in the US, UK, Australia, Canada and Sweden during 1981 – 2009. They use a flexible framework to estimate forward-looking monetary policy reaction functions with time-varying parameters. Taylor-type policy rules are augmented with the IMF financial stress index that is composed of three sub-indexes capturing several types of financial stress: banking stress, stock market stress and exchange rate stress. In general, the results show that financial stress is of little relevance to central banks when good financial conditions prevail. On the contrary, the monetary policy reaction to financial conditions seems to change substantially in times of high financial stress. In these episodes, policy rate changes can be explained to a large extent by the impact of financial instability, most evidently during the financial crisis in 2007 - 2009. For instance, the calculated financial stress effect on the interest rate implies that central banks set their policy rates around 50 - 100 basis points lower due to financial stress around the recent crisis. With respect to the components of the financial stress index, most central banks appear to put more emphasis on banking and stock market stress, while exchange rate stress is more relevant for policymakers in open economies, such as Canada and Sweden. Finally, Baxa, Horvath and Vasicek (2013) also briefly comment on the time-varying response parameters to inflation and output gap. Regarding the US, they demonstrate that the response to inflation has

[32] The weight is the probability of a financial crisis and is modelled as a function of the financial stress index exceeding its threshold value.

somewhat declined in the recent decade and even turned negative around the time of the financial crisis, while the response to output gap remained relatively stable.[33]

Alcidi, Flamini and Fracasso (2011) argue that credit spreads indicating the overall health of financial system may have an impact on how monetary policy is conducted in the US during the Greenspan era (1987:Q3 - 2005:Q4). They employ a logistic smooth transition regression model to estimate the Taylor rule augmented with the yield spread between the Moody's BAA corporate bond index and 10-year Treasury bonds, i.e. the credit spread. In order to model smooth transition across regimes, two variables are used: the credit spread to proxy for general concerns of policymakers regarding the health of financial system and the lagged policy rate to reflect the possibility of hitting the zero lower bound. The policy rule differs substantially between the high- and low-spread regimes. In the low-spread regime, all variables in the augmented Taylor rule are statistically significant and correctly signed. Meanwhile, it is only the lagged interest rate that determines monetary policy decisions in the high-spread regimes. With regards to the ZLB, it is demonstrated that after the policy rate falls below the 3% threshold, the reaction function of the Fed changes and the lagged interest rate term remains the only significant determinant in the augmented Taylor-type rule.

Similarly, Gnabo and Moccero (2015) employ a smooth transition regression and provide some evidence of the indirect Fed’s response to financial stress. They estimate a non-linear Taylor rule over the period 1987:Q4 – 2005:Q4. The transition between two regimes is modelled on the basis of the level of economic risk. The economic risk is captured using two measures: the measure of dispersion associated with the outlook of inflation derived from surveys and the VXO index by the Chicago Board Options Exchange as a proxy of financial market stress. The results show that the Fed responds to inflation and output gap positively and significantly in normal times. However, it becomes more responsive to output gap when the level of economic risk is higher, while the reaction to inflation does not seem to change between the regimes. On the other hand, the inflation parameter becomes statistically insignificant.

Overall, empirical evidence indicates that the Fed’s monetary policy framework tends to vary across the different regimes of financial conditions. The rest of the chapter examines the Fed’s monetary policy with respect to financial market developments.

[33] Vasicek (2012) tests for non-linearity with respect to financial stress in monetary policy rules in the Czech Republic, Hungary and Poland for the sample period 1998:1 – 2010:3. The results from the threshold regression of the augmented Taylor-type rule provide support for asymmetric policy behaviour. For instance, central banks in Czech Republic and Poland reduce their respective policy rates in light of elevated financial instability; however, the direct response to the financial stress index is mostly insignificant or incorrectly signed if the index value does not exceed the estimated threshold.

1.4 Methodology

As it has been discussed earlier, the Fed may be responding to financial market developments and this response is likely to be asymmetric. High financial stress periods tend to be accompanied with aggressive expansionary monetary policy; however, the Fed has not appeared to be eager to tighten policy stance in light of excessive asset price and credit booms in the past. Nevertheless, the literature offers rather mixed empirical evidence with respect to leaning against the wind in the US. Using a simple framework, this chapter provides the comprehensive analysis of the Fed's response to financial markets with respect to four different dimensions of financial stress related to the credit risk, stock market illiquidity, stock market conditions (bear versus bull markets) and overall financial conditions. Firstly, the direct reaction to financial markets is analysed and it is tested whether the response is asymmetric, i.e. varying across the business cycle. In addition, the analysis also considers the possibility of the indirect reaction to financial stress through changes in the response parameters on standard macroeconomic variables in the Taylor rule. The focus here is not solely on the global financial crisis as other episodes of financial turmoil over the sample period are also taken into account. Furthermore, the impact of the latest crisis is isolated in an effort to examine its relative importance for the overall findings.

Following Clarida, Gali and Gertler (1998), the regression analysis is based on a forward-looking Taylor rule with an interest rate smoothing term:

$$i_t = \left(1 - \sum_{j=1}^{n} \rho_j\right)\left(\alpha + \beta\pi_{t+k} + \gamma\hat{y}_t\right) + \sum_{j=1}^{n} \rho_j i_{t-j} + \varepsilon_t \tag{1.3}$$

where $\rho \in [0;1]$ denotes the degree of policy inertia (n = 2), π_{t+k} is an expected inflation rate k quarters ahead (k = 4), $\hat{y}_t$ represents contemporaneous output gap, and ε_t is an error term.[34] Equation (1.3) embeds the assumption that a central bank sets its policy rate with respect to expected inflation rate four periods ahead and current output gap. For the "Taylor principle" to hold, the inflation parameter is expected to be above unity ($\beta > 1$).

[34] Levin, Wieland and Williams (2003) demonstrate that it is the optimal choice to respond to one-year-ahead inflation forecasts and current-period output gap. Also, other studies also estimate Taylor rule specifications involving expected inflation and contemporaneous output gap (Hoffmann, 2013; Gnabo and Moccero, 2015).

With respect to output stabilisation, the parameter γ is expected to be positive implying a higher policy rate when output is above its long-term trend.

Several points should be kept in mind when estimating such a monetary policy rule. Typically, contemporaneous information on target macroeconomic variables is not available to policymakers at the time of decision making. Furthermore, there may exist reverse causality between the interest rate and explanatory variables. In order to address these issues, a great number of empirical studies utilise the Generalised Method of Moments (GMM) to estimate forward-looking Taylor-type rules (Clarida et al., 1998, 2000; Chadha et al, 2004; Fuhrer and Tootell, 2008; Mehra and Sawhney, 2010). Assuming rational expectations, the GMM framework allows replacing the unobserved (contemporaneous or forecast) values of independent variables with their actual (realised) values using some instruments. These instrumental variables should be exogenous with respect to the interest rate and should help forecast variables of interest, i.e. inflation and output gap, as well as be available to policymakers at the time of decision making (Clarida et al. 1998).[35]

There are several advantages of this methodology. Firstly, it helps to deal with the issues of unobserved data in real time as well as endogeneity. Secondly, it is relatively simple and easy to implement. Moreover, the GMM estimator is more robust with respect to a wide range of data generating process, i.e. the distribution of error terms, as compared to the full-information maximum likelihood estimators (Hansen, 1982; Baum, Shaffer and Stillman, 2003).

Nevertheless, there are also several disadvantages of using the GMM. For instance, it is well known that small-sample GMM estimates may be severely biased (Baum, Shaffer and Stillman, 2003). In addition to this, estimation results may depend on the choice of the estimator itself, i.e. the two-step GMM versus the continuously-updated GMM or the iterative GMM. Furthermore, the estimates also appear to depend on the procedure to calculate the optimal weighting matrix required in the GMM estimations (Jondeau, Bihan and Galles, 2004). Another drawback of this methodology is associated with instruments used. There is little theoretical motivation provided for the choice of instruments in the related literature and it is very challenging to find good instrumental variables (Siklos, Bohl and Werner, 2004; Nikolsko-Rzhevskyy, 2011). Variables can be considered as good instruments if they are exogenous with respect to the policy rate and they are highly correlated with endogenous regressors. However, it is very likely that instruments typically employed in the Taylor rule estimation may be only moderately correlated with

[35] The technical details of this method are provided in the Appendix B.

endogenous variables, i.e. instruments are weak, leading to biased GMM estimates (Baum and Shaffer, 2003; Mavroeidis, 2004). Also, it may not be appropriate to use the realised values of variables in the reaction function since they may not be "the cause" of the policy decisions but rather "the result" (Nikolsko-Rzhevskyy, 2011). Finally, Orphanides (2001) demonstrates that instrumental variables method using GMM with *ex post* data can easily obscure the fact that monetary policy is forward-looking.

As an alternative to the GMM, some studies use the maximum likelihood estimator (ML) (Gerlach-Kristen, 2004; Jondeau, Bihan and Galles, 2004). Its main advantage in the forward-looking context as compared to the GMM is that the expectations obtained with this methodology are fully consistent. Nevertheless, Jondeau, Bihan and Galles (2004) demonstrate that the ML and GMM estimates are very similar in the sample period starting after 1987. Following the seminal work by Orphanides (2001), it is increasingly popular to estimate the Taylor rule using real-time data for contemporaneous variables or real-time survey data on expectations about inflation and output. In this case, there is no need to instrument for forecasts of macroeconomic variables since they are formed using contemporaneously available data in real-time. Thus, the Ordinary Least Square estimator (OLS) can be used. A vast amount of more recent literature choose to employ the OLS with real–time forecast data (Mehra and Minton, 2007; Molodtsova, Nikolsko-Rzhevskyy and Papell, 2008; Orphanides and Wieland, 2008; Nikolsko-Rzhevskyy, 2011).

Consequently, the empirical analysis here is based on the OLS estimations using survey data on one-year-ahead inflation expectations and the *ex post* measure of current output gap.[36] Initially, Equation (1.3) is estimated as the benchmark policy rule. Next, it is augmented by including four different financial indicators (one by one) in order to test for the direct Fed's reaction to financial market developments:[37]

$$i_t = \left(1 - \sum_{j=1}^{n} \rho_j\right)\left(\alpha + \beta\pi_{t+k} + \gamma\hat{y}_t + \mu x_t\right) + \sum_{j=1}^{n} \rho_j i_{t-j} + \varepsilon_t \qquad (1.4)$$

where x_t is a selected financial variable at time t and μ represents a contemporaneous central bank's reaction to the financial indicator.

[36] In robustness analysis, the real-time data on real Gross Domestic Product (GDP) is used to construct the output gap. The main findings hold.

[37] With respect to financial variables, it is more difficult to defend the exogeneity assumption as it is possible that asset prices respond to monetary policy actions within the same period. Therefore, the robustness analysis is conducted using the GMM and reported in Section 1.7.1. The findings are qualitatively similar to the main results in Section 1.6. Hayford and Malliaris (2005) also employ the OLS for the main analysis, while providing the GMM estimates for the comparison.

The analysis considers four types of financial variables: a measure of overall financial conditions, an interest rate spread between risky and relatively safe long-term assets, i.e. the credit spread, stock market returns and a stock market liquidity measure. If the Fed responds to financial market developments, it is expected that the estimated coefficient is statistically significant and implies a lower policy rate in the periods of deteriorating financial conditions, and vice versa.[38]

As it has been previously argued, financial instability is likely to lead to sharp contraction in real economic activity. Hence, central bank's reaction to financial markets may be more pronounced in recessionary periods as compared to economic expansions, especially, if the cause of a recession originates from the financial system. In order to test for the Fed's response to financial markets across the business cycle, i.e. whether the response is asymmetric, a recession dummy variable is added to the estimated policy rule:

$$i_t = \left(1 - \sum_{j=1}^{n} \rho_j\right)\left(\alpha + \beta\pi_{t+k} + \gamma\hat{y}_t + \mu^D D^R x_t + \mu^{ND}\left(1 - D^R\right)x_t\right) + \sum_{j=1}^{n} \rho_j i_{t-j} + \varepsilon_t \tag{1.5}$$

where D^R is the dummy variable that takes the value of one to indicate US recessions as classified by the National Bureau of Economic Research (NBER), and zero otherwise. The reaction coefficients μ^D and μ^{ND} denote the Fed's response to a financial variable in economic recessions and expansions, respectively. It is expected that the monetary policy reaction to financial market developments is stronger in recessionary periods than in good economic conditions.

Several recent empirical studies indicate that in financial distress standard target variables could be of much less importance to policymakers and their response to them could change substantially (Alcidi, Flamini and Fracasso, 2011; Belke and Klose, 2013). Thus, it is worthwhile to examine whether the policy response to standard macroeconomic variables depends on financial market conditions, i.e. whether the Fed reacts to financial markets indirectly. In other words, it is tested how the Fed's reaction to expected inflation and output gap differ in the periods of high versus low financial stress.[39] Following Borio

[38] Whether μ is expected to be positive or negative depends on a selected variable as it is explained in the data section.

[39] With respect to non-linear monetary policy rules, several methods may be used to allow for regimes in the estimated rules: smooth transition regression models (Florio, 2009; Alcidi, Flamini and Fracasso, 2011; Castro, 2011; Martin and Milas, 2013), Markov-Switching models (Valente, 2003; Assenmacher-Wesche, 2006), and time-varying parameter regressions (Kim and Nelson, 2006; Baxa, Horvath and Vasicek, 2013). Also, other studies split the sample to account for the changes in regression parameters between two periods, for example, crisis versus pre-crisis (Belke and Klose, 2010), or use dummy variables to model regime switches (Borio and Lowe, 2004).

and Lowe (2004), a dummy-variable approach is employed to distinguish between the periods of severe financial stress and the periods of relatively favourable financial conditions. Dummy variables denoting financial distress are interacted with the response coefficients to inflation and output gap.

Firstly, appropriate thresholds are set with respect to each selected indicator of financial conditions and the financial stress dummy variables are constructed that take the value of one if the respective threshold is breached, indicating the periods of intense financial market distress, and zero otherwise. It is assumed that changes in the coefficients induced by financial stress occur suddenly and in a discrete manner, i.e. regime changes are determined by analysing the historical data on financial indicators and the past periods of high uncertainty in financial markets. Thus, the approach taken in this chapter is more closely related to the regime-switching methodology as opposed to smooth transition regression models. However, in regime-switching models different regimes are not identified *ex ante* but rather are estimated from the data and changes between regimes occur with a certain probability.

Secondly, the specification in Equation (1.3) is augmented using these dummy variables one at a time:

$$i_t = \left(1 - \sum_{j=1}^{n} \rho_j\right)\left(\alpha + \beta^D D^X \pi_{t+k} + \beta^{ND}\left(1 - D^X\right)\pi_{t+k} + \gamma^D D^X \hat{y}_t + \gamma^{ND}\left(1 - D^X\right)\hat{y}_t\right) + \sum_{j=1}^{n} \rho_j i_{t-j} + \varepsilon_t \quad (1.6)$$

where D^X is the financial stress dummy constructed on the financial variable x_t. The reaction coefficients to inflation and output gap during intense financial stress are denoted by β^D and γ^D, respectively. The estimated parameters are expected to be smaller in the periods of financial instability than in good financial conditions, i.e. $\beta^D < \beta^{ND}$ and $\gamma^D < \gamma^{ND}$.

1.5 Data and sample period

1.5.1 Sample period

The main empirical analysis is based on quarterly data for the US over the sample period that spans the Great Moderation (GM) and includes the global financial crisis. The significant decline in macroeconomic volatility since the mid-1980s motivates the choice of the beginning of the sample in 1985:Q1. In response to the financial crisis, the federal funds rate target was set to almost zero in December 2008. It has remained at exceptionally low levels for an extended period of time complicating the estimation of interest rate-based monetary policy rules. Following the adoption of unconventional policy tools, the funds target rate has become an inadequate proxy for monetary policy stance. Consequently, as the specifications of the Taylor rule in this chapter do not account for the ZLB, the sample period ends in 2008:Q4.

1.5.2 Data and variables

The effective federal funds rate (average of daily data) is used as a proxy for monetary policy rate (i_t). The measure of inflation expectations (π_{t+4}) is one-year-ahead annual inflation forecasts based on the GDP price index and provided by the Survey of Professional Forecasters (SPF). The output gap ($\hat{y}_t$) is defined as the percentage deviation of the seasonally adjusted log real GDP series from its Hodrick-Prescott trend. Data on the effective FFR and real GDP is obtained from FRED database maintained by the Federal Reserve Bank of St. Louis. The SPF data is provided by the Federal Reserve Bank of Philadelphia.

With respect to financial indicators (x_t), the Citi Financial Conditions Index (*CFCI*) is used as a proxy for broad financial conditions. The index includes information on corporate spreads, money supply, equity values, mortgage rates, the real trade-weighted dollar index, and energy prices. It is stated in terms of standard deviations from norms and the zero value is consistent with a normal expected pace of economic expansion. The positive values of the index indicate that financial variables are collectively exerting an expansionary force on the economy. Equivalently, negative index readings represent contractionary conditions. The credit spread is calculated as the difference between the Moody's BAA and AAA corporate bond yields (*CSPR*). Stock market returns are defined as the annual difference in the log S&P500 stock price index (*SP*).[40] In order to gauge stock market liquidity (*LQ*), the aggregate liquidity factor constructed by Pastor and Stambaugh (2003) is used (average of monthly data). The liquidity factor is calculated as the equally weighted average of the liquidity measures of individual stocks on the NYSE and AMEX, using daily data within the month. On average, the value of the market

[40] Annual asset price inflation is also used in other related studies (Belke and Klose, 2010; Belke and Klose, 2013).

liquidity measure is negative and more negative values indicate lower aggregate liquidity. The CFCI series is made available by Mark W. Watson.[41] The data on the Moody's corporate bond indices is taken from FRED database, while the S&P500 stock market index series is obtained from Datastream. Finally, the liquidity factor data is available from Lubos Pastor.[42] In terms of the sign of the estimated μ , it is expected to be positive for the CFCI, liquidity factor and stock market returns. For the credit spread, the response coefficient should be negative. Thus, in response to worsening financial conditions, lower aggregate stock market liquidity, declining stock prices and increasing credit spread, the policy rate is expected to be cut.

To set the threshold values that capture elevated stress levels and to construct the financial market stress dummy variables, the historical data on four financial indicators described above is used. All dummy variables take value 1 when financial stress is high and zero otherwise. The first dummy D^{CFCI} takes value 1 when the CFCI is below its historical average, i.e. the value of index is negative. Similarly, the credit spread dummy (D^{CSPR}) takes value 1 when the credit spread is above its historical average. The stock market distress is represented by the dummy (D^{SP}) that takes value 1 when the S&P500 index is below its 2-year moving average. Finally, D^{LQ} identifies financial stress when the aggregate stock market liquidity measure is below its historical average less one standard deviation.[43]

Figure 1.2 plots the financial variables together with their respective dummy variables. The shaded areas represent the NBER recession dates. It can be noted that the past episodes of financial stress, such as the stock market crash in 1987, are captured quite well. Also, the periods of heightened financial instability tend to coincide with US recessions. The stock market and CFCI stress dummies capture all three recessionary episodes, while the last recession, which commenced in December 2007, was associated with elevated levels in all four measures of financial stress.

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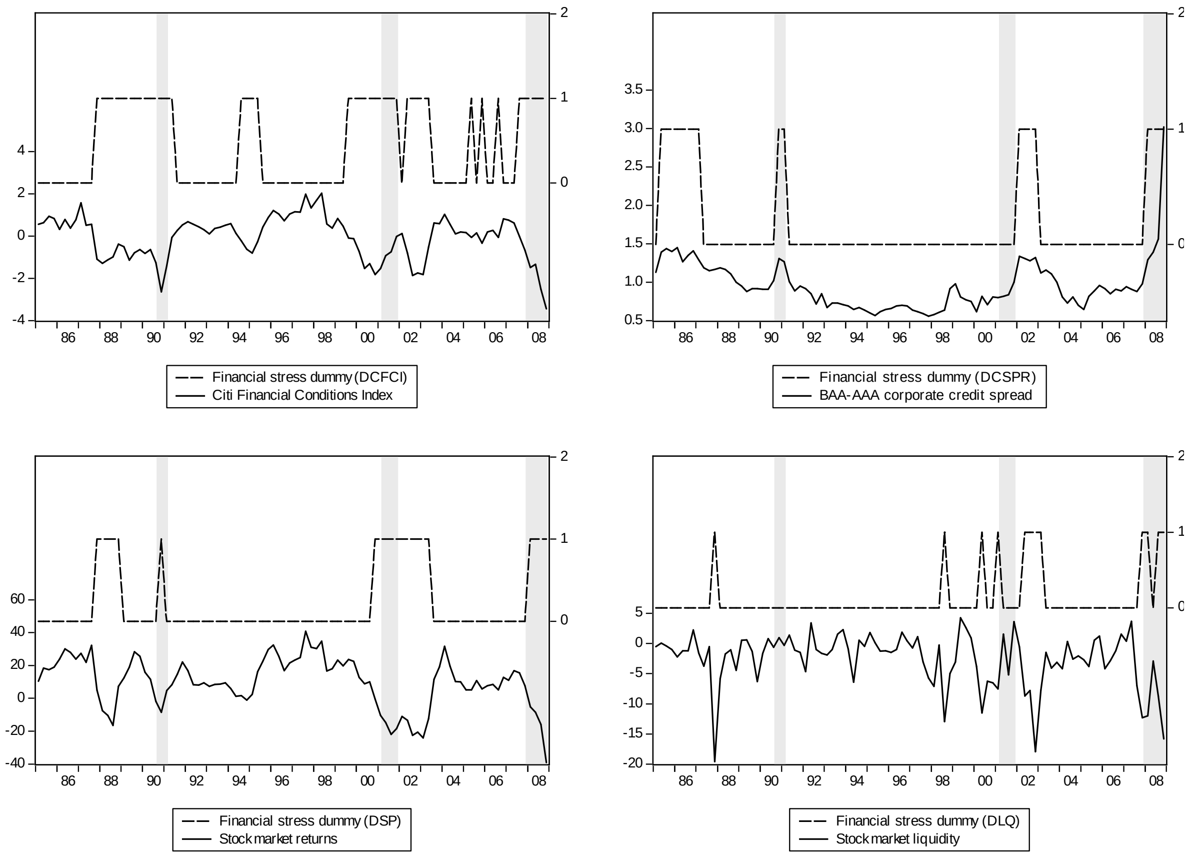
[41] The data can be accessed at http://www.princeton.edu/~mwatson/. More details on the construction of the index are provided in the appendix in Diclemente and Schoenholtz (2008).
[42] The time series for this liquidity measure is available at http://faculty.chicagobooth.edu/lubos.pastor/research/.
[43] The historical averages are calculated over the following periods: 1983:Q1 – 2009:Q4 (Citi FCI), 1919:Q1 – 2012:Q1 (credit spread), 1963:Q1 – 2011:Q4 (liquidity factor). The reason for subtracting one standard deviation from the liquidity factor average is that the financial distress indicator becomes too noisy if simply a historical average is used as the threshold.

Figure 1.2: Financial indicators and financial stress dummy variables



Notes: This figure plots four financial indicators with their respective financial stress dummy variables over the sample period 1985:Q1 – 2008:Q4. The top left panel presents the Citi Financial Conditions Index (*CFCI*) and the dummy variable D^{CFCI} that takes value 1 when the index is below its historical average. The top right panel presents the credit spread between Moody's BAA and AAA corporate bond yields (*CSPR*) and the dummy variable D^{CSPR} that takes value 1 when the credit spread is above its historical average. The bottom left panel presents annual stock returns on the S&P500 (*SP*) and the dummy variable D^{SP} that takes value 1 when the S&P500 index is below its 2-year moving average. The bottom right panel presents the stock market liquidity measure and the dummy variable D^{SP} that takes value 1 when liquidity measure is below its historical average less one standard deviation. The shaded areas denote the periods of US recessions as defined by the NBER.

1.6 Empirical results

1.6.1 Direct reaction to financial markets

To begin with, the basic Taylor rule in Equation (1.3) is estimated and the results are reported in Panel A of Table 1.1. The sum of two lagged interest rate terms is 0.89 indicating a high degree of policy inertia. The Taylor principle is not violated as the estimated response coefficient on expected inflation is well above one. In response to an increase in expected inflation rate by 1% (percentage point), a nominal short-term interest rate is raised by 2.13% that is enough to increase a real interest rate. Also, the coefficient is statistically significant at 1% level. The estimated parameter on the output gap is 0.98 and it is also highly statistically significant, implying a strong Fed's response to real economic activity. For instance, given an increase in output gap by 1 percentage point, the federal funds rate is raised by almost 1%. Thus, the Fed conducts countercyclical monetary policy. These findings are consistent with existing empirical studies (Clarida et al., 2000; Alcidi, Flamini and Fracasso, 2011; Nikolsko-Rzhevskyy, 2011).

Table 1.1: Basic and augmented Taylor rules – direct reaction to financial markets

	Panel A:	**Panel B: Augmented Taylor rules**			
	Basic Taylor rule	*CFCI*	*CSPR*	*SP*	*LQ*
α	-1.288	-1.315	1.608	-1.246	-0.176
	(1.504)	*(1.038)*	***(1.378)***	*(0.978)*	*(1.125)*
ρ_1	1.471***	1.442***	1.349***	1.400***	1.466***
	(0.106)	*(0.090)*	***(0.091)***	*(0.098)*	*(0.094)*
ρ_2	-0.579***	-0.568***	-0.477***	-0.539***	-0.579***
	(0.088)	*(0.079)*	***(0.078)***	*(0.085)*	*(0.083)*
β	2.127***	2.167***	2.341***	1.957***	1.90***
	(0.504)	*(0.374)*	***(0.436)***	*(0.344)*	*(0.409)*
γ	0.982**	1.074***	0.912**	0.845***	1.038***
	(0.491)	*(0.340)*	***(0.410)***	*(0.310)*	*(0.374)*
μ	-	1.027**	-3.667***	0.050**	0.186*
		(0.395)	***(1.207)***	*(0.022)*	*(0.111)*
Eq. SE	0.359	0.336	0.330	0.347	0.348
$\bar{R}^2$	0.974	0.977	0.978	0.976	0.976

Notes: This table reports the OLS estimates of Equation (1.3) in Panel A and Equation (1.4) in Panel B over the sample period 1985:Q1 – 2008:Q4. The financial indicators included into the augmented Taylor rule one by one are: the Citi Financial Conditions Index (*CFCI*), the credit spread between Moody's BAA and AAA corporate bonds (*CRSP*), annual stock returns on the S&P500 index (*SP*), and the stock market liquidity measure by Pastor and Stambaugh (2003) (*LQ*). Standard errors are reported in parentheses. Appropriate standard errors are used based on the White heteroscedasticity test and Ljung-Box Q-statistics. White heteroscedasticity-consistent standard errors are reported in *italic*, while heteroscedasticity and autocorrelation-consistent (HAC) standard errors are reported in ***bold italic***. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\bar{R}^2$ denotes adjusted R-squared.

In order to examine the direct policy response to financial market developments, the augmented Taylor rule in Equation (1.4) is estimated either with the financial conditions index, credit spread, annual stock market returns or aggregate liquidity measure as an additional variable. Panel B of Table 1.1 summarises the results. Firstly, the degree of interest rate smoothing remains very similar across alternative specifications as compared to the basic rule estimation. Secondly, the estimated response parameter on inflation is around 2 and is always statistically significant. The parameter β is the smallest (1.90) in the case of the Taylor rule augmented with the liquidity factor and the largest (2.34) when the credit spread is included. The reaction coefficient on output gap remains statistically significant at conventional levels with the magnitude of around 1. It is the largest (1.07) for the specification with the financial conditions index and the smallest (0.85) when annual stock price enters the policy rule.

With respect to the financial variables, the estimated parameters for the CFCI, stock returns and credit spread are statistically significant at either 1% or 5% levels. Meanwhile, the response to the liquidity factor is found to be only marginally significant. Thus, the liquidity conditions in the stock market appear to be of less importance to the Fed's policymakers. Following a decline in the financial conditions index, that implies deteriorating financial conditions, the federal funds target rate is reduced. Thus, monetary policy stance is more expansionary when financial conditions are tight. Worsening financial conditions may eventually lead to the contraction in economic activity and lower price level; thus, policymakers may wish to respond to declining financial conditions index with a rate cut as it that lowers the costs of borrowing, increases the value of collateral and other assets as well as boosts confidence among consumers and financial market participants. Similarly, the Fed also reduces its policy rate in response to a decline in aggregate liquidity, falling stock prices and to a higher credit spread which all are associated with detiorating conditions in financial markets. For instance, an increase in the credit spread by 1% results in the federal funds rate that is lower by 3.67% than otherwise. This indicates a very strong policy response to credit market conditions. In comparison to the baseline Taylor rule, the standard errors of regression are somewhat smaller in the case of augmented Taylor rules, indicating a better fit once financial market developments are allowed to enter the reaction function of the Fed.

The result showing the strong policy response to financial conditions is consistent with the findings in Montagnoli and Napolitano (2005). Also, many others find that the Fed conducts expansionary policy when stock prices decline (Borio and Lowe, 2004; Chadha, Sarno and Valente, 2004). The estimated μ for the credit spread (-3.67) is similar to the estimated coefficients in Castelnuevo (2003) and Gerlach-Kristen (2004).

Overall, the results summarised in Table 1.1 suggest that the Fed responds directly to financial market developments in addition to the response to standard macroeconomic variables. Consequently, the policy rate is typically lower in times of financial stress and it is higher when financial conditions are good. Nevertheless, anecdotal historical evidence impies that the Fed is likely to treat changes in financial indicators differently depending on existing financial conditions and business cycle (Roubini, 2006). Thus, the policy reaction function may be asymmetric.

1.6.2 Direct reaction to financial markets: recession vs. expansion

This section compares the direct Fed's response to financial indicators in the periods of economic recession versus economic expansion. It is expected that monetary policymakers pay more attention to financial market developments when setting the policy rate in recessions as compared to normal economic conditions. Table 1.2 presents the estimation results of Equation (1.5).

Table 1.2: Augmented Taylor rules – direct reaction to financial markets across the business cycle

	CFCI	*CSPR*	*SP*	*LQ*
α	-0.794 *(0.986)*	0.255 *(1.012)*	-0.331 *(1.040)*	0.064 *(1.058)*
ρ_1	1.380*** *(0.081)*	1.319*** *(0.072)*	1.378*** *(0.088)*	1.385*** *(0.096)*
ρ_2	-0.499*** *(0.069)*	-0.442*** *(0.059)*	-0.497*** *(0.076)*	-0.511*** *(0.085)*
β	2.098*** *(0.379)*	1.982*** *(0.331)*	1.839*** *(0.380)*	1.752*** *(0.371)*
γ	1.027*** *(0.291)*	1.034*** *(0.252)*	0.872** *(0.366)*	1.215*** *(0.270)*
μ^{ND}	0.194 *(0.252)*	-0.678 *(0.828)*	0.012 *(0.024)*	0.029 *(0.051)*
μ^{D}	2.600** *(1.056)*	-4.336*** *(1.087)*	0.236** *(0.104)*	0.523*** *(0.179)*
Eq. SE	0.311	0.285	0.327	0.319
$\overline{R}^2$	0.981	0.984	0.979	0.980
$\boldsymbol{WT^X}$	0.033	0.004	0.060	0.005

Notes: This table reports the OLS estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator in economic expansions and economic recessions, respectively. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The last row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^{D}$. See also Table 1.1 notes.

With respect to interest rate smoothing and the response to standard macroeconomic variables, the results are consistent with Table 1.1. The Fed sets its policy rate in response to both inflation and output gap and allows for a slow adjustment of the interest rate towards its target as indicated by the significant lagged interest rate terms. The estimate of β ranges between 1.75 and 2.10 across the specifications, thus, always being well above unity. With respect to the output gap parameter, it varies from 0.87 to 1.22.

In line with expectations, there is strong evidence of the asymmetric policy reaction to financial variables. The response coefficients (μ^{ND}) are significantly smaller in magnitude during economic expansions as compared to recessionary periods (μ^{D}). Moreover, the estimates are also statistically insignificant in good times. In contrast, the Fed's reaction to financial indicators is strong and highly statistically significant in recessionary periods for all specifications. For instance, in response to a one-standard-deviation decrease in the financial conditions index, the Fed would cut the policy rate target by 2.6%.[44] Similarly, the policy rate would be decreased by 4.34% if the credit spread increases by 1%. Meanwhile, a 10% drop in annual stock market returns would lead to a policy rate cut by 2.36%. Finally, if the liquidity cost of trading stocks increases by 1% (on a $1m trade in 1962 stock market dollars), the Fed responds to the declining liquidity factor by lowering the policy rate by 0.52%.[45]

According to the Wald test, the null hypothesis that $\mu^{D} = \mu^{ND}$ can be rejected at 5% level in the case of the CFCI, credit spread and liquidity factor and at 10% level in the case of stock returns. As compared to Table 1.1 results, the response parameters in recessionary periods are also larger in magnitude than the estimates of μ with respect to all financial variables. This implies aggressive and economically significant monetary policy easing in response to adverse financial developments during economic recessions, while no such response is evident during economic expansions. Furthermore, allowing for the asymmetric response to financial markets appears to describe the actual policy behaviour better since the regression standard errors are now smaller than for the symmetric specifications in Table 1.1.

The above findings are in line with the study by Kasai and Naraidoo (2012) who demonstrate that monetary policymakers in South Africa respond to financial conditions more strongly during recessions than expansions.[46] As shown in Figure 1.2, recessionary periods are usually associated with declining stock returns and worsening overall financial conditions. Thus, the results reported here are also consistent with the existing evidence that the Fed tends to respond to financial markets only when financial conditions deteriorate. For instance, Ravn (2012) finds that the Fed eases policy stance when stock prices decline, but it does not tighten when stock prices rise. Similarly, Mattesini and Becchetti (2009) show that the Fed uses expansionary policy to support excessively undervalued stocks, while it does not tend to curb the positive deviations of asset prices from their fundamental values by increasing the policy rate. In addition, Baxa, Horvath and Vasicek (2013) find that the effect of financial conditions on monetary policy decisions is

[44] The index is stated in terms of standard deviations from the mean value.
[45] The liquidity factor becomes more negative (decreases).
[46] They estimate non-linear augmented Taylor rules using a smooth transition regression model with output as a transition variable. Nevertheless, this asymmetry disappears during the global financial crisis.

strong during the periods of intense financial stress implying a lower policy rate than otherwise. On the other hand, the effect is not found to be present in normal times.

1.6.3 Indirect reaction to financial markets: high vs. low financial stress

To examine whether the Fed's policy framework changes in times of elevated financial market stress, the reaction coefficients on inflation and output gap in estimated Taylor rules are allowed to vary across the periods of high and low financial stress. The estimation results of Equation (1.6) are presented in Table 1.3.

Table 1.3: Basic Taylor rules - indirect reaction to financial markets

	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}
α	-1.648 *(1.340)*	-1.639 ***(1.616)***	-1.261 ***(1.452)***	-0.478 *(1.106)*
ρ_1	1.490*** *(0.092)*	1.425*** ***(0.079)***	1.446*** ***(0.087)***	1.456*** *(0.098)*
ρ_2	-0.597*** *(0.081)*	-0.533*** ***(0.068)***	-0.561*** ***(0.073)***	-0.569*** *(0.086)*
β^{ND}	2.486*** *(0.531)*	2.377*** ***(0.589)***	2.218*** ***(0.517)***	1.928*** *(0.411)*
β^{D}	2.000*** *(0.453)*	1.689*** ***(0.558)***	1.716*** ***(0.632)***	0.947 *(0.778)*
γ^{ND}	1.245** *(0.498)*	0.965** ***(0.440)***	1.026** ***(0.429)***	1.019*** *(0.380)*
γ^{D}	1.090* *(0.598)*	1.053 ***(1.224)***	0.753 ***(0.864)***	1.023 *(0.995)*
Eq. SE	0.355	0.353	0.356	0.353
$\bar{R}^2$	0.975	0.975	0.975	0.975
$\boldsymbol{WT^{\pi}}$	0.126	0.169	0.246	0.148
$\boldsymbol{WT^{\hat{y}}}$	0.837	0.941	0.737	0.997

Notes: This table reports the OLS estimates of Equation (1.6) over the sample period 1985:Q1 – 2008:Q4 using four financial stress dummy variables (D^X). D^{CFCI} indicates financial stress related to overall financial conditions, D^{CSPR} denotes credit risk-related stress, D^{SP} identifies stock market bear conditions and D^{LQ} denotes stock market liquidity-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of γ^{ND} and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The last two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table 1.1 notes.

When favourable financial conditions prevail, the interest rate response to inflation (β^{ND}) is statistically significant at 1% level and is always much greater than one. The largest coefficient of 2.49 is reported for the specification where financial stress is defined using the dummy constructed on the basis of the financial conditions index. The lowest value of 1.93 is obtained when financial stress is defined by the stock market liquidity. Hence, this indicates strong anti-inflationary preferences of monetary policymakers at the

Fed in good financial conditions. The response parameter to output gap (γ^{ND}) is also positive and statistically significant across the specifications with the estimated value ranging from 0.97 to 1.25. The Fed conducts countercyclical policy in times of no financial stress with the strong reaction to both macroeconomic goal variables. These findings are in line with the benchmark rule estimation in Table 1.1.

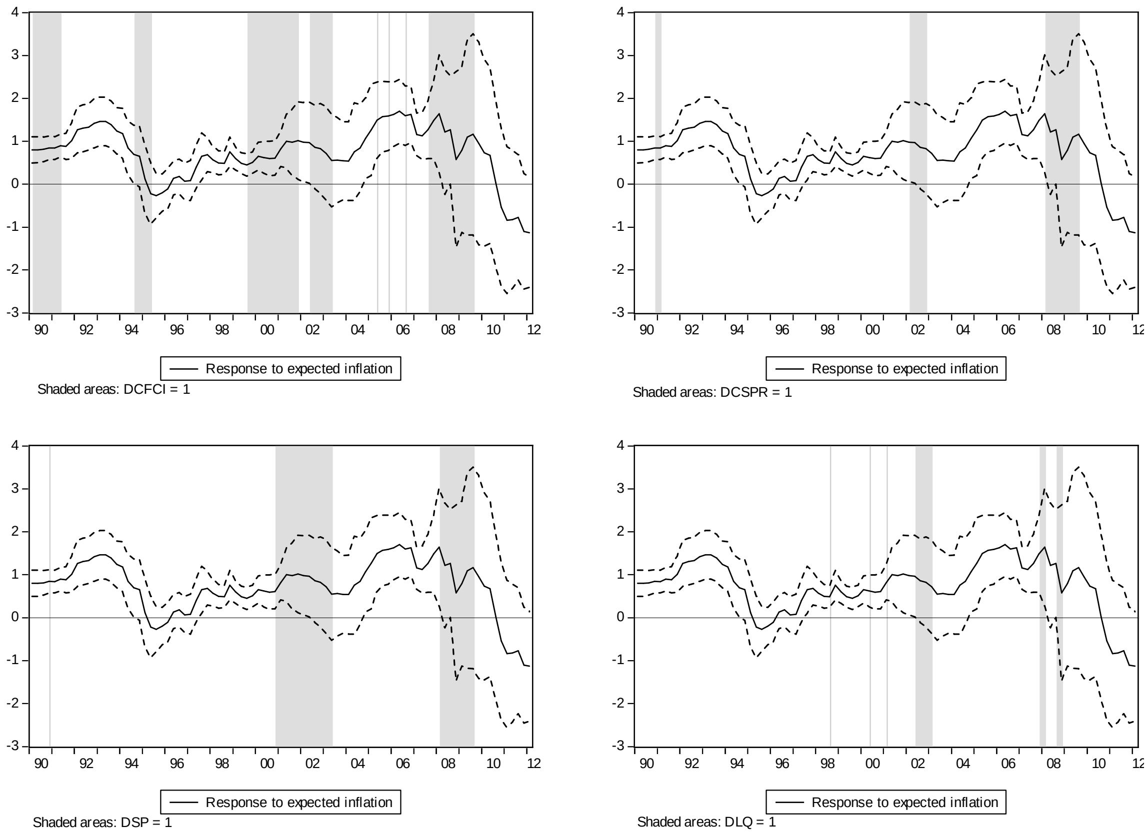
On the other hand, the estimates are somewhat different when financial conditions worsen considerably, i.e. when financial stress dummies become active. Firstly, the parameter β^{D} is smaller in magnitude than β^{ND} for all specifications, ranging from 0.95 in the case of stock market liquidity stress to 2.00 when overall financial conditions deteriorate. Furthermore, the estimated β^{D} violates the Taylor principle and is insignificant when aggregate liquidity in the stock market falls considerably. Nevertheless, according to the Wald test, the two inflation parameters are never significantly different. Secondly, the Fed's reaction to output gap becomes generally statistically insignificant in times of financial distress. However, the magnitude of the estimated output gap parameter remains broadly unchanged, lying between 0.75 and 1.09. As in the case of inflation, the difference between γ^{D} and γ^{ND} is statistically insignificant according to the Wald test. There appears to be only a moderate decline in the Fed's response to expected inflation during the periods of financial market stress and it typically remains highly significant and above unity. Also, the response parameter with respect to output gap does not appear to change overall; however, the response becomes statistically insignificant in times of bad financial conditions. Thus, the Fed still cares enough about inflation even if there are signs of substantial financial market stress, although there is some statistical uncertainty around its response to output gap.

In general, these findings do not provide strong evidence in favour of substantial changes in the Fed's policy framework with respect to macroeconomic conditions in times of financial stress as reported by other studies. For instance, Baxa, Horvath and Vasicek (2013) show that the Fed's response to inflation declined substantially following the terrorist attacks in 2001 and during the recent crisis, whilst the parameter on output gap increased but only slightly. On the other hand, Gnabo and Moccero (2015) do not find that the Fed's reaction to inflation changes in the periods of heightened financial risk as measured by expected stock market volatility. Nevertheless, the response parameter on output gap is found to increase considerably during financial uncertainty. The difference in the results for the inflation parameter across the two studies may be due to a shorter sample period used in the latter as they do not consider the global financial crisis (Gnabo and

Moccero, 2015). Nevertheless, the findings in Table 1.3 with respect to output gap are consistent with Alcidi, Flamini and Fracasso (2011). They also find that the output gap coefficient becomes statistically insignificant in the periods of high credit spreads.

In addition, Figures 1.3 and 1.4 provide some insight into how the monetary policy response to inflation and output gap, respectively, varies over time. The figures plot the regression coefficients that are obtained by estimating the basic Taylor rule over a five-year rolling window together with 95% confidence bands.[47] The shaded areas represent the periods of high financial stress as defined by a respective dummy variable.

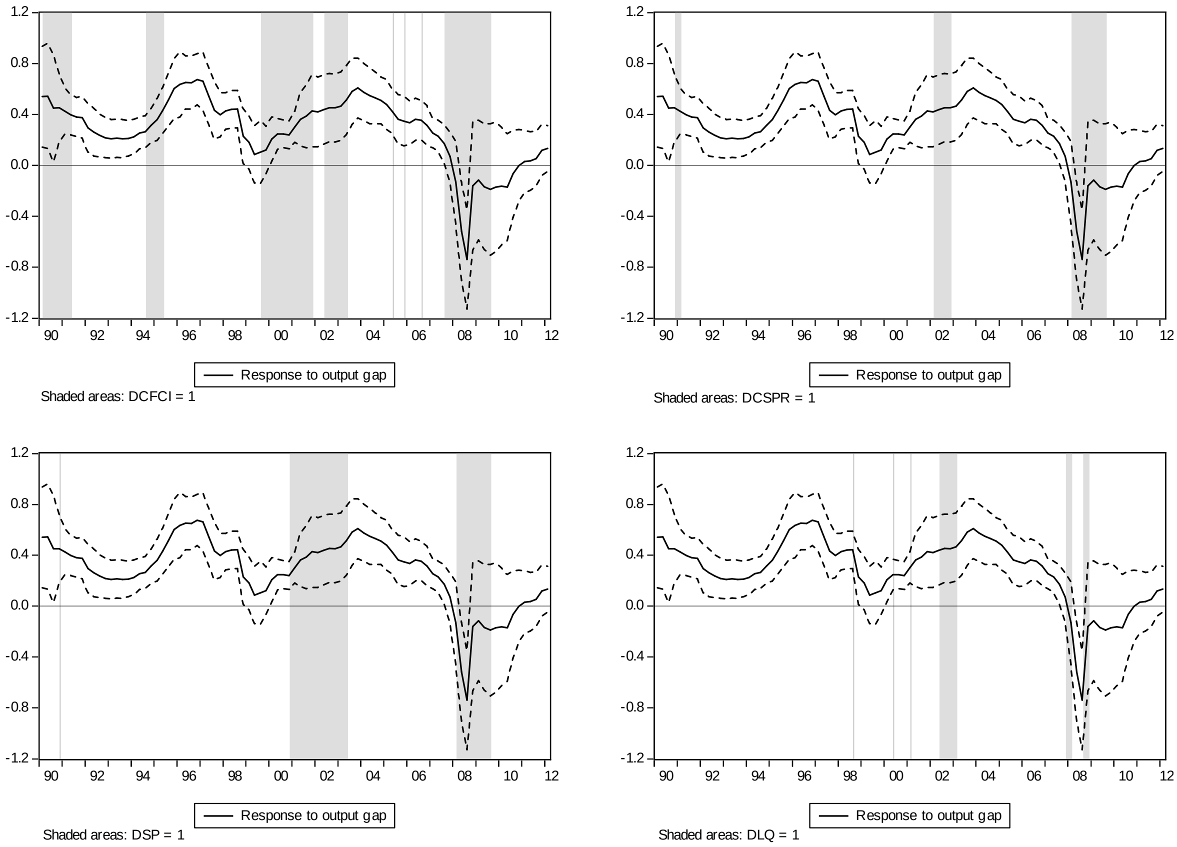
Figure 1.3: The estimated response to expected inflation



Notes: This figure plots the response coefficient to expected inflation (solid line) together with +/- 2 standard error bands (dashed lines) obtained from the five-year rolling-window OLS estimation of an equation: $i_t = \alpha + \beta\pi_{t+4} + \gamma\hat{y}_t + \rho i_{t-1} + \varepsilon_t$. The first estimation window spans 1985:Q1 – 1990:Q1. The shaded areas denote the periods of intense financial stress as indicated by a respective financial stress dummy: overall financial stress (top left panel, D^{CFCI} =1), credit risk-related stress (top right panel, D^{CSPR} =1), stock market stress (bottom left panel, D^{SP} =1) and stock market liquidity-related stress (bottom right panel, D^{LQ} =1).

[47] The plotted coefficients are obtained from the rolling five-year window OLS estimations of an equation $i_t = \alpha + \beta\pi_{t+4} + \gamma\hat{y}_t + \rho i_{t-1} + \varepsilon_t$. The initial window spans 1985:Q1 – 1990:Q1 and the sample ends at 2012:Q1. The sample extends beyond the end of 2008 as it allows seeing more clearly the dynamics of coefficients around the crisis period.

Figure 1.4: The estimated response to output gap



Notes: This figure plots the response coefficient to output gap (solid line) together with +/- 2 standard error bands (dashed lines) obtained from the five-year rolling-window OLS estimation of an equation: $i_t = \alpha + \beta\pi_{t+4} + \gamma\hat{y}_t + \rho i_{t-1} + \varepsilon_t$. The first estimation window spans 1985:Q1 – 1990:Q1. The shaded areas denote the periods of intense financial stress as indicated by a respective financial stress dummy: overall financial stress (top left panel, D^{CFCI} =1), credit risk-related stress (top right panel, D^{CSPR} =1), stock market stress (bottom left panel, D^{SP} =1) and stock market liquidity-related stress (bottom right panel, D^{LQ} =1).

In line with the estimation results discussed earlier, the reaction coefficient on inflation typically declines in times of intense financial stress as shown in Figure 1.3. With respect to output gap, the response parameter appears to be somewhat larger in the periods of financial distress; however, this pattern does not seem to hold during the latest crisis. During the recent financial crisis both parameters fell sharply and became negative and statistically insignificant.

The finding of negative response coefficients to policy goal variables is counter-intuitive and implies that the Fed decreases the funds rate target if inflation and output gap increase. Nevertheless, this does not necessarily reflect such systematic behaviour of policymakers. The finding of negative parameters with respect to inflation and output gap may be explained by the events around the time of the financial crisis, i.e. since the mid-2007. Firstly, the expected one-year-ahead inflation, as reported by the SPF, remained stable and there were no signs by the end of 2008 that inflation expectations declined sharply in response to the financial crisis. Secondly, the output gap was positive and increasing through 2007, although it declined afterwards and finally turned negative only in the final quarter of 2008. In the meantime, the federal funds rate target was cut repeatedly starting in September 2007 through December 2008. After the funds rate hit the zero lower bound, the crisis eventually led to lower expected inflation and negative output gap values, while at the same time the interest rate remained at the near-zero level. This could explain why there is an indication of a negative relationship between the interest rate and macroeconomic variables. Furthermore, the pressure in financial markets was increasing sharply since 2007 reaching the peak in September 2008. It is likely that central bankers started to pay relatively more attention to adverse financial conditions that prevailed at that time. Thus, the Fed was potentially more responsive to financial conditions rather than to changes in expected inflation and/or output gap during the crisis period, leading to the structural break in the estimated Taylor rule.

These sharp changes in the reaction coefficients around the financial crisis are not reflected in the estimates provided in Table 1.3. The crisis effect may be obscured as the sample period also includes other episodes of financial turmoil rather than just the most recent one. Hence, the past episodes of financial stress appear to dampen the impact of the events in 2007 – 2008.

1.6.4 The effect of financial crisis in 2007 - 2008

In order to separately evaluate the impact of the global financial crisis, a new dummy variable D^{07-08} is constructed that only takes the value of 1 during 2007:Q4 – 2008:4Q, and zero otherwise. The dummy variables used in Equations (1.5) and (1.6) are replaced with the new dummy and the results are reported in Table 1.4.

The first column of Table 1.4 presents the estimates from the standard Taylor rule where the weights on inflation and output gap are allowed to change. The results indicate a major change in the Fed’s response to inflation. Specifically, the inflation parameter decreases substantially from 1.73 prior to the crisis to -1.41 during the crisis, and also turns statistically insignificant in the crisis period. The Wald test identifies a significant structural change in the estimated inflation parameter. The negativity of β^D may be reflecting the fact that the Fed cut the funds target rate in response to severe financial stress despite inflation expectations not changing substantially, i.e. remaining relatively stable, until the end of 2008. Also, this implies that the Fed potentially “ignored” the expected inflation variable and instead was more concerned about the conditions in financial markets at that time, i.e. policymakers acted pre-emptively based on financial data since inflation expectations data had not yet reflected signs of considerably worsening outlook. This result is in line with recent evidence by Baxa, Horvath and Vasicek (2013) for the US and Martin and Milas (2013) for the UK. The policy response to output gap remains highly significant over the period 2007-2008 and appears to even increase in magnitude (from 1.26 to 1.99), albeit not sufficiently for the Wald test to identify a significant shift as compared to the pre-crisis period. Hence, the Fed continues to respond strongly to the output gap measure, indicating strong preferences for economic growth. As compared to the estimated parameters in Table 1.3, the inclusion of the earlier episodes of financial stress, which were less severe, seems to attenuate the effect of the global financial crisis on the estimates of the indirect Fed's reaction to financial variables.

Table 1.4: Basic and augmented Taylor rules – financial crisis effect

	Panel A: Basic Taylor rule	**Panel B: Augmented Taylor rules**			
	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*
α	0.062 ***(0.866)***	-0.465 *(0.800)*	0.490 ***(1.238)***	-0.780 *(0.908)*	0.091 *(0.900)*
ρ_1	1.293*** ***(0.084)***	1.361*** *(0.078)*	1.325*** ***(0.072)***	1.381*** *(0.092)*	1.367*** *(0.093)*
ρ_2	-0.445*** ***(0.068)***	-0.504*** *(0.069)*	-0.468*** ***(0.063)***	-0.523*** *(0.080)*	-0.503*** *(0.079)*
β	-	1.911*** *(0.295)*	1.893*** ***(0.314)***	1.898*** *(0.330)*	1.743*** *(0.321)*
γ	-	1.090*** *(0.242)*	1.069*** ***(0.277)***	0.821*** *(0.301)*	1.196*** *(0.243)*
μ^{ND}	-	0.484** *(0.205)*	-0.956 ***(0.983)***	0.031* *(0.017)*	0.028 *(0.051)*
μ^{D}	-	2.475** *(0.944)*	-3.765*** ***(1.053)***	0.209*** *(0.071)*	0.563*** *(0.152)*
β^{ND}	1.730*** ***(0.299)***	-	-	-	-
β^{D}	-1.411 ***(0.881)***	-	-	-	-
γ^{ND}	1.260*** ***(0.251)***	-	-	-	-
γ^{D}	1.990*** ***(0.521)***	-	-	-	-
Eq. SE	0.311	0.318	0.311	0.332	0.312
$\bar{R}^2$	0.981	0.980	0.981	0.978	0.980
$\boldsymbol{WT^X}$	-	0.037	0.006	0.019	0.001
$\boldsymbol{WT^\pi}$	0.000	-	-	-	-
$\boldsymbol{WT^{\hat{y}}}$	0.162	-	-	-	-

Notes: This table reports the OLS estimates of Equation (1.6) in Panel A and Equation (1.5) in Panel B over the sample period 1985:Q1 – 2008:Q4 using financial crisis dummy variable (D^{07-08}). D^{07-08} takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator before the crisis and during the crisis, respectively. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation before the crisis and during the crisis, respectively. Estimates of $\gamma^{ND}$and γ^{D} denote policy response to output gap before the crisis and during the crisis, respectively. The last three rows (WT^X, WT^π and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^D$, $\beta^{ND} = \beta^D$ and $\gamma^{ND} = \gamma^D$, respectively. See also Table 1.1 notes.

The remaining columns of Table 1.4 summarise the results from the estimations of Equation (1.5) with the new crisis dummy. The findings imply that the financial crisis had the considerable impact on the direct response by the Fed to the financial indicators. Specifically, it is only the CFCI variable that is significant at a conventional significance level in the period prior to the crisis, albeit the coefficient μ^{ND} is notably lower as compared to the corresponding financial crisis estimate μ^D. In contrast, all four financial market variables are significant at 1% or 5% levels in the estimated augmented Taylor rule during the recent crisis. Furthermore, the response parameter μ^D is significantly different from μ^{ND} with respect to each financial indicator. For instance, prior the crisis 1% (percentage point) increase in the credit spread does not reduce the policy rate. During the financial crisis, the same change in the spread implies a 377-basis-point decrease in the federal funds rate. In addition, the sign and magnitude of the coefficients is typically quite close to the respective estimates in economic recessions reported in Table 1.2. Thus, the evidence of the direct response to financial markets identified in Tables 1.1-1.2 appears to be largely driven by financial developments and the Fed actions since late 2007.

To sum up, the main empirical analysis implies several things about the policy of the Federal Reserve. Firstly, it seems to be fairly well described by a forward-looking Taylor rule allowing for interest rate smoothing. Secondly, the Fed has reacted to financial market developments in addition to information about expected inflation and output. Nevertheless, the direct response to financial indicators is found to be highly asymmetric and dependent on the business cycle conditions. During economic recessions, monetary policy is eased much more aggressively in response to deteriorating financial conditions. On the other hand, this results appear to be driven to a great extent by the period of the global financial crisis. With respect to the indirect reaction to financial market developments, there seems to be only a moderate decline in the Fed' response to expected inflation in times of intense financial stress. Meanwhile, there is no significant change in the reaction to output gap, albeit the output gap parameter turns insignificant. However, if the financial crisis of 2007-2008 is separately taken into account, the evidence of the indirect response to financial markets strengthens considerably. The parameter on expected inflation declines significantly and even turns negative in light of financial distress, implying a significant impact of the crisis on the Fed's monetary policy conduct. It appears that the Fed typically follows a standard Taylor-type policy rule; however, in times of intense financial stress, such as the global financial crisis, financial market developments also become important in setting the policy rate, while the role of inflation potentially decreases.

It is important to note that the analysis in this chapter does not attempt to answer the question of whether central bankers should or should not respond to financial market developments. It merely provides an insight as to whether they do and, if they do, how they respond to financial information.

1.7 Robustness analysis

The robustness of the main findings in Section 1.6 is tested in several ways. Firstly, the GMM method is employed to estimate the monetary policy rule. Secondly, alternative financial variables are included into the Taylor rule. Thirdly, the reaction functions in Section 1.6 are re-estimated using the real-time data on real GDP. Furthermore, an alternative measure of output gap is used, i.e. the potential output is calculated by applying a quadratic trend on real GDP series. Finally, the estimations are carried out for alternative sample periods. The main results are found to be reasonably robust to all changes. The results are reported in the Appendix A.

1.7.1 Estimations using GMM

Due to the potential simultaneous interaction between the policy interest rate and independent variables in the Taylor rule, the OLS estimates may be biased. While monetary policymakers set their policy rate in response to inflation expectations and output gap, it is also likely that macroeconomic variables are contemporaneously affected by interest rate decisions. It could be argued that "sluggish" macroeconomic variables, such as inflation and output gap, are slow to respond to monetary policy shocks and it may take more than a month to adjust (Christiano, Eichenbaum and Evans, 2005). On the other hand, monetary policy could have some impact on macroeconomic variables over the period of one quarter, i.e. the frequency used in the analysis of this chapter. With respect to inflation, real-time survey data on inflation expectations is used, thus, the endogeneity should not be an issue in this case.[48] With respect to output gap, the simultaneity bias is more likely as *ex post* data is used.

[48] Note that Survey of Profesional Forecasters typically reports this expectations data in the middle month of a respective quarter. Thus, if a policy decision is made at the beginning of a quarter, such data is not yet available to policymakers.

Furthermore, the problem of endogeneity becomes even more severe when financial variables enter the policy rule. On the one hand, central banks may have reasons to respond to developments in asset prices such as stock prices as well as spreads between interest rates on risky and safe assets as explained in Section 1.3. On the other hand, asset prices contain all information available to market participants and any new information, including monetary policy news, is immediately priced in. For instance, monetary policy rate changes can have an impact on stock prices through its effects on expected future cash flows and the discount rate used to discount these cash flows (Bernanke and Kuttner, 2005; Bjornland and Jacobsen, 2013; Kontonikas and Kostakis, 2013). In addition, some studies find that monetary policy may also have significant effects on credit spreads (Cenesizoglu and Essid, 2012). Given that the financial conditions index includes the information on asset prices and bond yields among other variables, it is also likely to be influenced by monetary policy shocks within the same quarter quarter. Furthermore, both policymakers and financial markets could be responding simultaneously to some macroeconomic news, leading to the biased estimates of the policy response to financial market developments.

To this regard, the endogeneity issue is addressed by estimating Equations (1.3)-(1.6) using the GMM that is commonly employed in the Taylor-rule literature (Clarida et al., 1998, 2000; Chadha, Sarno and Valente, 2004).[49] Following Belke and Klose (2013), the HAC (Newey-West) weighting matrix that provides the heteroscedasticity and autocorrelation consistent estimator of the long-run covariance matrix is chosen.[50] The theoretical framework and technical details of this methodology are presented in the Appendix B.

The set of instruments may consist of any lagged variables that are useful to forecast endogenous regressors, such as inflation, output and financial variables, in addition to any contemporaneous variables that are exogenous with respect to the interest rate. In line with the great majority of the literature, the instruments used in the Taylor rule specifications without any financial variables are a constant and four lags (t = -1 to t = -4) of federal funds rate, expected one-year-ahead inflation, output gap and 10-year Treasury yield (Clarida et al., 1998; 2000; Mehra and Sawhney, 2010; Castro, 2011). These variables are potentially good predictors of future inflation and economic activity. Thus, in total there are seventeen instruments in specifications without a financial variable. With respect to the augmented Taylor rules, four lags of a respective financial variable are added in addition to the instruments listed above. This is a standard approach in the literature

[49] The results also hold if the lagged (by one quarter) values of financial variables are included in the estimated augmented Taylor rules using the GMM.

[50] A Bartlett kernel with Newey-West bandwidth selection is used.

since asset prices and measures of financial conditions are likely to influence future price and output levels (Chadha et al., 2004; Castro, 2011, Belke and Klose, 2013; Hofmann, 2013). Also, the past values of asset prices may be good predictors of future asset prices and financial conditions. Thus, the number of instruments in these specifications increase to twenty one.

The results are provided in Tables A1.1-A1.4. As it can be noted from Table A1.1, the basic rule estimates are in line with Table 1.1. Both the inflation and output gap parameters are positive and highly significant. The Taylor principle holds. All financial indicators in the augmented specifications are statistically significant with correctly signed coefficients. For each specification, the p-value of the *J*-statistics implies that the model is well-specified and that over-identifying restrictions cannot be rejected. In line with the main findings, Table A1.2 provides the evidence of the asymmetric policy reaction to financial indicators. As shown previously, the Fed’s response to all financial variables is significantly greater during economic recessions, while there is no evidence that additional variables have any significant impact on setting the policy rate during expansionary periods. The parameters on inflation and output gap remain largely unchanged.

With respect to the indirect response to financial markets, the estimation results are reported in Table A1.3. Across alternative specifications, the reaction coefficient on inflation declines in times of intense financial stress. In the case of stock market distress and liquidity-related financial stress, the decline in β^{D} is statistically significant (see the Wald test p-values). With respect to the remaining financial stress dummies, the difference between β^{ND} and β^{D} is not found to be significant. The response to output gap is higher when financial conditions deteriorate; however, it is usually insignificant regardless of the state of financial markets. In line with the results in Table 1.3, the parameters γ^{ND} and γ^{D} do not appear to be significantly different according to the Wald test. Finally, Table A1.4 provides some insight into the global financial crisis impact on the Fed’s policy. The findings are overall in line with the results reported in Table 1.4. During the crisis, the response parameter on inflation decreases sharply and turns negative with the Wald test indicating that this change is statistically significant. Meanwhile, there appears to be no substantial change with respect to the output gap coefficient. The Fed’s response to financial indicators is found to be significant during the financial crisis and all reaction coefficients increase significantly in magnitude as compared to the pre-crisis period. Furthermore, there is no indication of significant effects from any financial variable on the policy rate before the crisis.

The related literature typically focuses on the *J*-statistics alone to infer about the validity of the model. However, this does not imply anything about the weakness of instruments used. Therefore, Tables A1.1 – A1.4 report the Kleibergen-Paap rk Wald *F*-statistics for each specification to test for the presence of weak instruments, i.e. when instrumental variables are correlated with endogenous regressors but only weakly (Baum, Shaffer and Stillman, 2007). The null hypothesis is that the equation is weakly identified and the rejection of this null indicates the absence of weak instruments problem. As robust GMM options are used, the critical values calculated by Stock and Yogo (2005) cannot be used to test whether the null can be rejected. Thus, the rule of thumb is applied that the *F*-statistic should be at least as large as 10 in order to reject the null (Staiger and Stock, 1997). In addition, a partial R^2 measure by Shea (1997) is presented for each endogenous regressor to provide some insight into the relevance of instruments. Essentially, it is a measure of the correlation between instruments and endogenous explanatory variables that accounts for intercorrelations among instruments. Overall, there is some indication of the presence of weak instruments, especially as indicated by the Kleibergen-Paap *F*-statistic. Nevertheless, in the majority of cases the partial R^2 measure is relatively large for all endogenous regressors, i.e. it is above 0.50.

1.7.2 Alternative financial variables

In this section, four alternative financial indicators are included one by one into Equation (1.4): the Chicago National Financial Conditions Index (*NFCI*), the credit spread between the Moody's Baa corporate bonds and 10-year US Treasury bonds (*CSPR2*), the Macroeconomic Advisers Monetary and Financial Conditions Index (*MAFCI*), and the interbank spread between the 3-month LIBOR and 3-month US Treasury bill rates (*IBSPR*).[51] The positive readings of the NFCI represent financial conditions that are tighter than on average and the negative values of the index imply that financial conditions are looser than on average. Thus, the estimated μ is expected to be negative. On the other hand, negative values of the MAFCI indicate financial conditions that lead to a contraction in real GDP, while positive values show expansionary financial conditions. Therefore, the parameter on the MAFCI is expected to be positive. With respect to the credit and interbank spreads, μ should be negative.

[51] Data series for the NFCI, 10-year Treasury yield (constant maturity), 3-month Treasury bill rate and 3-months LIBOR rate are obtained from the FRED database. The MAFCI data is obtained at http://www.princeton.edu/~mwatson/.

As previously, each financial variable is used to construct a financial stress dummy variable that takes value 1 when the financial stress is high and zero otherwise. The first dummy D^{NFCI} takes value 1 when the NFCI is above its historical average, i.e. when the values are positive. Similarly, the credit spread dummy (D^{CSPR2}) takes value 1 when the credit spread is above its historical average. With respect to the MAFCI, the financial stress is represented by the dummy D^{MAFCI} that takes value 1 when the index reading is negative.[52] Finally, D^{IBSPR} identifies financial stress when the spread is above its historical average.[53]

The results are presented in Tables A1.5 – A1.8. The response coefficients to inflation and output gap are in line with expectations. The evidence of the direct Fed's response to financial variables in Table A1.5 is somewhat mixed. Two variables, namely the NFCI and credit spread, are statistically significant and the parameters have the expected (negative) sign. On the other hand, the interbank spread and MAFCI are not found to be significant determinants of the policy interest rate. In line with the main results, Table A1.6 indicates that the Fed's reaction to financial market developments is asymmetric. In recessionary periods, all financial variables enter the Taylor rule with a highly statistically significant parameter that has the expected sign. On the contrary, during economic expansions it is only the reaction to the credit spread that remains statistically significant, albeit it is weaker. Furthermore, the null hypothesis of $\mu^{ND} = \mu^{D}$ cannot be rejected only in the case of the NFCI.

The findings summarised in Tables A1.7-A1.8 are broadly in line with the main results in Tables 1.3-1.4. The reaction coefficient on expected inflation declines moderately in the periods of intense financial stress but remains significant. The Wald test indicates that the difference between two inflation coefficients is typically statistically insignificant. The evidence on the response to output gap is rather mixed. There is also little indication that γ^{ND} is statistically different from γ^{D}. The results reported in Table A1.8 indicate that the financial crisis period is the key driver of the Fed's reaction to financial market developments. In the pre-crisis period, it is only the credit spread that is statistically significant in the augmented Taylor rule. On the other hand, the policy response to financial indicators is significantly stronger and typically statistically significant at 1% level during the crisis.

[52] With respect to the CFCI and NFCI their historical averages are equal to zero, thus, zero is also used as a threshold value for the MAFCI.

[53] Historical averages are calculated over the following sample periods: 1973:Q1 – 2012:Q1 (NFCI), 1962:Q1 – 2012:Q1 (credit spread), and 1986:Q1 – 2012:Q1 (interbank spread).

1.7.3 Real-time data

Instead of the *ex post* revised data on real GDP, the real-time data is used to construct an alternative measure of output gap. The real-time data is obtained from the Real-Time Data Set for Macroeconomists provided by the Federal Reserve Bank of Philadelphia.[54] The starting point is to apply the HP filter separately on the log real GDP series across all quarterly data vintages over the period 1985:Q1 - 2008:Q4. This way, the estimate of the potential real GDP is obtained for each quarter using only the historical data available in that quarter.[55] Next, for each quarter output gap is calculated by subtracting the final estimate of the potential log real GDP from the last available log real GDP value in the respective data vintage. Both the real-time and *ex post* measures of output gap are plotted in Figure A1.1. The differences are quite substantial in some periods, thus, the results may differ depending on which type of data is used for the estimation of policy rules.

The results are reported in Tables A1.9-A1.12. With respect to the basic Taylor rule, the response parameters on inflation and output gap are positive and statistically significant at 1% level. The Taylor principle is never violated. However, the estimated direct reaction of the Fed to developments in financial markets appears to be rather weak. As shown in Table A1.9, only the credit spread parameter is statistically significant. With respect to the asymmetric policy reaction, the estimates in Table A1.10 are in line with the main findings discussed in Section 1.6. During recessionary periods, financial indicators typically have a significantly stronger impact on the interest rate as compared to expansionary periods, albeit the parameter on the CFCI is only marginally significant. In line with the main findings, the Fed does not appear to consider financial market developments when setting the policy rate in good economic conditions.

From Table A1.11 it can be noted that the estimated response to inflation tends to decrease during intense financial stress, but it remains significant and is always greater than one. Meanwhile, the reaction to output gap typically increases, but turns statistically insignificant in the periods of financial distress. However, the parameters on both variables are not significantly different across two regimes as indicated by the Wald test. This is broadly consistent with the results reported in Table 1.3. Finally, Table A1.12 shows the impact of the financial crisis on the Fed's reaction function. During the crisis, the Fed's

[54] Inflation expectations measure is not changed because the SPF one-year-ahead inflation forecast is a real-time measure.

[55] Each data vintage contains data that goes back to 1947:Q1. The last entry point is the measure of the output in the previous quarter. For instance, the last available real GDP measure in the data vintage as of 1999:Q4 refers to 1999:Q3.

response to inflation drops sharply and turns insignificant, albeit the Wald test does not indicate a significant change. The coefficient γ^{D} increases in size considerably, but it is also insignificant and not significantly different from γ^{ND}. Finally, the reaction to financial indicators appears to be significant only during the crisis. The estimated μ^{ND} is significantly smaller in magnitude and is insignificant for all four financial variables. Thus, the evidence of significant changes in response to inflation or output gap due to financial stress is somewhat weaker as compared to the main results. On the other hand, the findings with respect to the asymmetric response to financial indicators are qualitatively very similar to the results in Section 1.6.

1.7.4 Alternative output gap measure

In this section, the quadratic trend is applied to the log real GDP series in order to construct an alternative measure of output gap. As can be noted from Tables A1.13-A1.14, the response to expected inflation and output gap is strong, positive and always significant. There is also the evidence of the significant Fed's reaction to financial market developments with most of financial indicators being significant in the augmented Taylor rules at conventional significance levels. In line with the main findings, this reaction is highly asymmetric with respect to the business cycle. During economic recessions, the Fed response to financial markets increases significantly as compared to economic expansions. Moreover, there is no evidence of the significant reaction to financial indicators in good economic times.

Table A1.15 presents the results from the estimated Equation (1.6). As previously found, the inflation parameter is always lower in times of financial market stress, albeit the Wald test does not indicate a significant difference between β^{ND} and β^{D} in most of specifications. The response to output gap typically becomes insignificant when financial conditions deteriorate; however, there is no indication of significant differences between two coefficients across the regimes. As shown in Table A1.16, the response to inflation becomes statistically insignificant during the financial crisis. The parameter on inflation declines, while the output gap parameter increases. The null of $\beta^{ND} = \beta^{D}$ can be rejected at 1% level, while the null of $\gamma^{ND} = \gamma^{D}$ cannot be rejected, even though the response to output gap is much smaller prior to the crisis. The policy reaction to financial market developments increases during the most recent financial crisis; however, it is typically statistically significant even in the pre-crisis period. In the case of the liquidity factor, the estimated μ^{D} is significantly greater than μ^{ND}. Overall, the findings are reasonably in line with the main results in Tables 1.3-1.4.

1.7.5 Alternative sample periods

Firstly, the original sample period is extended to coincide with the beginning of the Volcker's era and then it is shortened to start with the Chairmanship of Greenspan. Therefore, two alternative samples are considered in this section using the data as described earlier: 1979:Q4 – 2008:Q4 and 1987:Q4 – 2008:Q4.[56] The results are provided in the Tables A1.17 – A1.20. Table A1.17 reports the estimations of the basic (Panel A) and augmented (Panel B) Taylor rules. The results are consistent with those provided in Table 1.1. The inflation parameter is statistically significant and it complies with the Taylor principle. The reaction coefficient on output gap is always positive and statistically significant. The CFCI, credit spread, and stock returns enter the estimated policy rule with statistically significant and correctly signed parameters, while the liquidity factor does not appear to play any significant role in the policy reaction function. In line with the main findings, the direct Fed's reaction to the financial variables is typically significantly stronger in recessionary periods. Furthermore, it is mostly insignificant during economic expansions.

Table A1.19 reports the estimation results from Equation (1.6) for both sample periods. Generally, the Fed's response to inflation is higher in times of favourable financial conditions and it declines in the periods of elevated financial stress. Nevertheless, in the majority of specifications the Wald test does not indicate a significant difference between β^{ND} and β^{D} at conventional levels of significance. Thus, there only seems to be a moderate change in the Fed's response to inflation due to high financial stress. In terms of reaction to output gap, the evidence is somewhat mixed. Overall, the response parameter is positive and mostly statistically significant in normal times, but it typically declines in magnitude and becomes insignificant during heightened financial instability. Again, the Wald test fails to provide support for any significant difference between γ^{ND} and γ^{D}. These findings are consistent with the main results in Section 1.6. Finally, the effect of the recent financial crisis is examined and shown in Table A1.20. As previously, there is a clear indication that the inflation parameter declines significantly during the crisis and

[56] With respect to the longer sample period, only one lag of the federal funds rate is used as the second lagged term is not significant. One exception is the CFCI as the data starts only in 1983 and two lags are used.

becomes statistically insignificant. There is no substantial change in the response to output gap; however, the estimate increases slightly in magnitude. With respect to financial indicators, the Fed's reaction increases during the financial crisis and, typically, the increase is statistically significant. Also, the reaction to financial indicators tends to be insignificant in the pre-crisis period. Thus, the finding of the strong direct response of the Fed to financial market developments appears to be explained to a large extent by the end of the sample. Overall, the main results in Tables 1.3 and 1.4 are robust to alternative sample periods.

1.8 Conclusion

The global financial crisis has re-opened the old debate about whether monetary policy authorities should or should not respond to financial imbalances that may have extremely adverse effects on real economic activity. Alongside, there has been a surge in empirical studies that attempt to answer the question whether financial indicators play any role in a central bank's reaction function. This chapter begins with the brief discussion of the origins and development of the Taylor rule as well as some practical issues encountered in the literature on monetary policy rules. The empirical analysis is focused on the link between monetary policy and financial market developments. While some studies argue that the Federal Reserve reacts to financial market developments, the evidence is somewhat mixed. Using data for the US covering the period 1985:Q1 - 2008:Q4 and employing the alternative specifications of a forward-looking augmented Taylor rule, the chapter re-examines this conjecture. The analysis provides a simple, but deep and comprehensive study of the Fed's behaviour employing a range of financial variables across the different dimensions of financial market stress. Two main questions are investigated. Firstly, does the Fed react directly to the indicators of financial stress and, if so, is such reaction symmetric? Secondly, does the policy response to inflation and output gap change in the presence of intense financial stress? These questions are examined with respect to the four different dimensions of financial market stress: credit risk, stock market liquidity risk, stock market bear conditions and overall financial conditions.

The results provide the empirical evidence in support of both the direct and indirect reaction of the Fed to financial markets. Nevertheless, this reaction appears to be largely driven by monetary policy behaviour during the financial crisis in 2007-2008. While stock market returns, the credit spread, the measure of stock market liquidity and the financial conditions index are found to be statistically significant in the augmented (symmetric)

Taylor rule, financial market developments tend to have a significant impact on the policy rate only in recessionary periods. Moreover, the significant reaction to financial indicators during economic recessions is, to a great extent, driven by the Fed's actions in response to the global financial crisis. With respect to the indirect response, the Fed's reaction to expected inflation declines moderately, while the output gap parameter typically becomes insignificant in the periods of elevated financial stress. The indirect response to financial market stress strengthens significantly during the global financial crisis, i.e. 2007:Q4 – 2008:Q4. The parameter on expected inflation declines significantly, turns negative and statistically insignificant. With respect to output gap, the estimated coefficient increases slightly, but not substantially, and remains significant.

Overall, the results imply a lower policy rate in times of severe financial market stress, especially in the period 2007-2008. As a result of the financial crisis, the Fed's policy framework changed significantly with less emphasis on the price stability and, possibly, more focus on financial market conditions. This view is in line with the evidence in Baxa, Horvath and Vasicek (2013) for the US and Martin and Milas (2013) for the UK.

References for Chapter 1

Ahmad, S., 2016. A multiple threshold analysis of the Fed's balancing act during the Great Moderation. *Economic Modelling*, 55, p.343-358.

Alcidi, C., Flamini, A., and Fracasso, A., 2011. Policy regime changes, judgment, and Taylor rules in the Greenspan era. *Economica*, 78(309), p.89-107.

Assenmacher-Wesche, K., 2006. Estimating central bank's preferences from a time-varying empirical reaction function. *European Economic Review*, 50(8), p.1951-1974.

Assenmacher-Wesche, K., and Gerlach, S., 2010. Monetary policy and financial imbalances: facts and fiction. *Economic Policy*, 25(63), p.437-482.

Batini, N., and Haldane, A. 1999. Forward-looking rules for monetary policy. In John B. Taylor, 1999. *Monetary Policy Rules*, London: University of Chicago Press. Ch. 4.

Bauducco, S., Bulir, A., and Cihak, M., 2008. Taylor rule under financial instability. *IMF Working Papers*, Working paper WP/08/18.

Baum, C.F., Shaffer, M.E., and Stillman, S., 2003. Intrumental variables and GMM: Estimation and testing. *The Stata Journal*, 3(1), p.1-31.

Baum, C.F., Shaffer, M.E., and Stillman, S., 2007. Enhanced routines for instrumental variables/GMM estimation and testing. *Boston College*, Economics working paper 677.

Baxa, J., Horvath, R., and Vasicek, B., 2013. Time-varying monetary-policy rules and financial stress: Does financial instability matter for monetary policy? *Journal of Financial Stability*, 9(1), p.117-138.

Bean, C., 2003. Asset prices, financial imbalances and monetary policy: are inflation targets enough?. *BIS Working Papers*, Working paper 140.

Bec, F., Salem, M. B., and Collard, F., 2002. Asymmetries in monetary policy reaction function: Evidence for U.S., French and German central banks. *Studies in Nonlinear Dynamics and Econometrics*, 6(2), Article 3, p.1-20.

Belke, A., and Klose, J., 2010. (How) Do the ECB and the Fed react to financial market uncertainty? The Taylor rule in times of crisis. *DIW Berlin*, Discussion paper 972.

Belke, A., and Klose, J., 2013. Modifying Taylor reaction functions in the presence of the zero-lower-bound — Evidence for the ECB and the Fed. *Economic Modelling*, 35, p.515-527.

Bernanke, B.S., and Gertler, M., 1999. Monetary policy and asset price volatility. *Federal Reserve Bank of Kansas City Economic Review*, 84(4), p.17-51.

Bernanke, B.S., and Gertler, M., 2001. Should central banks respond to movements in asset prices?. *The American Economic Review*, 91(2), p.253-257.

Bernanke, B.S., and Kuttner, B.N., 2005. What explains the stock market's reaction to Federal Reserve policy? *Journal of Finance*, 60(3), p.1221-1257.

Bjornland, H.C., and Jacobsen, D.H., 2013. House prices and stock prices: different roles in the US monetary transmission mechanism. *The Scandinavian Journal of Economics*, 115(4), p.1084-1106.

Bjornland, H.C., and Leitemo, K., 2009. Identifying the interdependence between US monetary policy and the stock market. *Journal of Monetary Economics*, 56(2), p.275-282.

Bleich, D., Fendel, R., and Rulke, J., 2013. Monetary policy and stock market volatility. *Deutsche Bundesbank*, Working paper 45/2013.

Boivin, J., 2006. Has U.S. monetary policy changed? Evidence from drifting coefficients and real-time data. *Journal of Money, Credit and Banking*, 38(5), p.1149-1173.

Borio, C., 2004. Market distress and vanishing liquidity: Anatomy and policy options. *Bank for International Settlements*, Working paper 158.

Borio, C., 2014. Monetary policy and financial stability: what role in prevention and recovery?. *BIS Working Papers*, Working paper 440.

Borio, C., and Lowe, P., 2002. Asset prices, financial and monetary stability: Exploring the nexus. *BIS Working Papers*, Working paper 114.

Borio, C., and Lowe, P., 2004. Securing sustainable price stability: Should credit come back from the wilderness?. *BIS Working Papers*, Working paper 157.

Botzen, W.J. Wouter, and Marey, P.S., 2010. Did the ECB respond to the stock market before the crisis?. *Journal of Policy Modeling*, 32(3), p.303-322.

Brave, S., and Butters, R.A., 2002. Diagnosing the financial system: Financial conditions and financial stress. *International Journal of Central Banking*, 8(2), p.191-239.

Bullard, J.B, and Schaling, E., 2002. Why the Fed should ignore the stock market. *Federal Reserve Bank of St. Louis Review*, 84(2), p.35-42.

Bunzel, H., and Enders, W., 2010. The Taylor rule and "opportunistic" monetary policy. *Journal of Money, Credit and Banking*, 42(5), p.931-949.

Carlson, M.A., Lewis, K.F., and Nelson, W.R., 2012. Using policy intervention to identify financial stress. *Federal Reserve Board Finance and Economics Discussions Series*, Working paper 2012-12.

Castelnuovo, E., 2003. Taylor rules, omitted variables, and interest rate smoothing in the US. *Economics Letters*, 81(1), p.55-59.

Castro, V., 2011. Can central banks' monetary policy be described by a linear (augmented) Taylor rule or by a nonlinear rule?. *Journal of Financial Stability*, 7(4), p.228-246.

Cecchetti, S.G., 2009. Crisis and responses: The Federal Reserve in the early stages of the financial crisis. *The Journal of Economic Perspectives*, 23(1), p.51-76.

Cecchetti, S.G., Genberg, H., Lipsky, J., and Wadhwani, S., 2000. Asset prices and central bank policy. *Centre for Economic Policy Research*, The Geneva report on the World Economy, 2.

Cenesizoglu, T., and Essid, B., 2012. The effect of monetary policy on credit spreads. *The Journal of Financial Research*, 35(4), p.581-613.

Chadha, J.S., Sarno, L., and Valente, G., 2004. Monetary policy rules, asset prices, and exchange rates. *IMF Staff Papers*, 51(3), p.529-552.

Christiano, L., Ilut, C.L., Motto, R., and Rostagno, M., 2010. Monetary policy and stock market booms. *NBER Working Papers Series,* Working paper 16402.

Christiano, L.J., Eichenbaum, M., and Evans, C., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy,* 113(1), p.1-45.

Chung, H., Laforte, J.P., Reifschneider, D., and Williams, J.C., 2012. Have we underestimated the likelihood and severity of zero lower bound events?. *Journal of Money, Credit and Banking*, 44(1), p.47-82.

Clarida, R., Gali, J., and Gertler, M., 1998. Monetary policy rules in practice: Some international evidence. *European Economic Review*, 42(6), p.1033-1067.

Clarida, R., Gali, J., and Gertler, M., 2000. Monetary policy rules and macroeconomic stability: Evidence and some theory. *The Quarterly Journal of Economics*, 115(1), p.147-180.

Cogley, T., 1999. Should the Fed take deliberate steps to deflate asset price bubbles?. *Federal Reserve Bank of San Francisco Economic Review*, 1, p.42-52.

Cukierman, A., and Gerlach, S., 2003. The inflation bias revisited: Theory and some international evidence. *The Manchester School*, 71(5), p.541-565.

Cukierman, A., and Muscatelli, A., 2008. Nonlinear Taylor rules and asymmetric preferences in central banking: Evidence from the United Kingdom and the United States. *The B.E. Journal of Macroeconomics*, 8(1), Article 7, p.1-29.

Curdia, V., and Woodford., 2010. Credit spreads and monetary policy. *Journal of Money, Credit and Banking*, 42(S1), p.3-35.

D'Amico, S., English, W., Lopez-Salido, D., and Nelson, E., 2012. The Federal Reserve's large-scale asset purchase programmes: Rationale and effects. *The Economic Journal* 122(564), p.415-446.

Davis, S., and Huang, K.X.D., 2013. Optimal monetary policy under financial sector risk. *Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute*, Working paper 85.

Diclemente, R., and Schoenholtz, K., 2008. A view of the US subrpime crisis. Citigroup Global Markets, Report (unpublished).

Dolado, J., Maria-Dolores, R., and Naveira, M., 2005. Are monetary-policy reaction functions asymmetric?: The role of nonlinearity in the Phillips curve. *European Economic Review*, 49(2), p.485-503.

Dolado, J., Maria-Dolores, R., and Ruge-Murcia, F.J., 2004. Nonlinear monetary policy rules: Some new evidence for the U.S. *Studies in Nonlinear Dynamics and Econometrics*, 8(3), Article 2, p.1-32.

English, W., Tsatsaronis, K., and Zoli, E., 2005. Assessing the predictive power of measures of financial conditions for macroeconomic variables. *BIS Papers*, 22, p.228-252.

Evans, C.L., 1998. Real-time Taylor rules and the federal funds futures market. *Federal Reserve Bank of Chicago Economic Perspectives*, 22(3), p.44-55.

Fawley, B.W., and Neely, C.J., 2013. Four stories of quantitative easing. *Federal Reserve Bank of St. Louis Review*, 95(1), p.51-88.

Ferguson, R.W.Jr., 2005. Recessions and recoveries associated with asset-price movements: what do we know?. Speech at Stanford Institute for Economic Policy Research, 12 January, 2005.

Filardo, A., 2004. Monetary policy and asset price bubbles: calibrating the monetary policy trade-offs. *BIS Working Papers*, Working paper 155.

Finocchiaro, D., and von Heideken, V.Q., 2013. Do central banks react to house prices?. *Journal of Money, Credit and Banking*, 45(8), p.1659-1683.

Florio, A., 2006. Asymmetric interest rate smoothing: The Fed approach. *Economics Letters*, 93(2), p.190-195.

Florio, A., 2009. Asymmetric preferences for interest rate variability and non-linear monetary policy inertia. Sc*ottish Journal of Political Economy*, 56(5),p.685-704.

Freixas, X., Martin, A., and Skeie, D., 2011. Bank liquidity, interbank markets, and monetary policy. *The Review of Financial Studies*, 24(8), p.2656-2692.

Friedman, M., 1968. The role of monetary policy. *The American Economic Review*, 58(1), p.1-17.

Fuhrer, J., and Tootell, G., 2008. Eyes on the prize: How did the Fed respond to the stock market?. *Journal of Monetary Economics*, 55(4), p.796-805.

Furlanetto, F., 2011. Does monetary policy react to asset prices? Some international evidence. *International Journal of Central Banking*, 7(3), p.91-111.

Gagnon, J., Raskin, M., Remache, J., and Sack, B., 2011. The financial market effects of the Federal Reserve's large-scale asset purchases. *International Journal of Central Banking*, 7(1), p.3-43.

Gambacorta, L., and Signoretti, F.M., 2014. Should monetary policy lean against the wind? An analysis based on a DSGE model with banking. *Journal of Economic Dynamics and Control*, 43, p.146-174.

Gelain, P., Lansing, K.J., and Mendicino, C., 2013. House prices, credit growth, and excess volatility: Implications for monetary and macroprudential policy. *International Journal of Central Banking*, 9(2), p.219-276.

Gerlach-Kristen, P., 2004. Interest-rate smoothing: Monetary policy inertia or unobserved variables?. *Contributions to Macroeconomics*, 4(1), Article 3, p.1-17.

Gilchrist, S., and Zakrajsek, E., 2011. Monetary policy and credit supply shocks. *IMF Economic Review*, 59(2), p.195-232.

Gnabo, J., and Moccero, D., 2015. Risk management, nonlinearity and aggressiveness in monetary policy: The case of the US Fed. *Journal of Banking and Finance*, 55, p.281-294.

Goodhart, C., and Hofmann, B., 2002. Asset prices and the conduct of monetary policy. Working paper (unpublished).

Hafer, R.W., and Haslag, J.H., 1988. The FOMC in 1987: The effects of a falling dollar and the stock market collapse. *Federal Reserve Bank of St. Louis Review*, 70(2), p.3-16.

Hakkio, C.S., and Kahn, G.A., 2014. Evaluating monetary policy at the zero lower bound. *Federal Reserve Bank of Kansas City Economic Review*, 99(2), p.5-32.

Hakkio, C.S., and Keeton, W.R., 2009. Financial stress: What is it, how can it be measured, and why does it matter?. *Federal Reserve Bank of Kansas City Economic Review*, 94(2), p.5-50.

Hansen, L.P, 1982. Large sample properties of Generalised Method of Moments estimators. *Econometrica*, 50(4), p.1029-1054.

Hatzius, J., Hooper, P., Mishkin, F.S., Schoenholtz, K.L., and Watson, M.W., 2010. Financial conditions indexes: A fresh look after the financial crisis. *NBER Working Paper Series*, Working paper 16150.

Hayashi, F., 2000. *Econometrics*. 1st ed. Princeton: Princeton University Press.

Hayford, M.D., and Malliaris, A.G., 2005. How did the Fed react to the 1990s stock market bubble? Evidence from an extended Taylor rule. *European Journal of Operational Research*, 163(1), p.20-29.

Hoffmann, A., 2013. Did the Fed and ECB react asymmetrically with respect to asset market developments?. *Journal of Policy Modeling*, 35(2), p.197-211.

Hofmann, B., and Bogdanova, B., 2012. Taylor rules and monetary policy: A global „Great Deviation“?. *BIS Quarterly Review*, September, p.37-49.

Hubrich, K., and Tetlow, J.R., 2015. Financial stress and economic dynamics: The transmision of crises. *Journal of Monetary Economics*, 70, p.100-115.

Jondeau, E., Bihan, H.L., and Galles, C., 2004. Assessing Generalised Method-of-Moments estimates of the Federal Reserve reaction function. *Journal of Business and Economic Statistics*, 22(2), p.225-239.

Jovanovic, M., and Zimmermann, T., 2010. Stock market uncertainty and monetary policy reaction functions of the Federal Reserve bank. *The B.E. Journal of Macroeconomics*, 10(1), Article 21, p.1-17.

Judd, J.P., and Rudebusch, G.D., 1998. Taylor's rule and the Fed: 1970–1997. *Federal Reserve Bank of San Francisco Economic Review*, 3, p.3-16.

Kahn, G.A., 2010. Taylor rule deviations and financial imbalances. *Federal Reserve Bank of Kansas City Economic Review*, 95(2), p.63-99.

Kasai, N., and Naraidoo, R., 2012. Financial assets, linear and nonlinear policy rules. An in-sample assessment of the reaction function of the South African Reserve Bank. *Journal of Economic Studies*, 39(2), p.161-177.

Kahn, G.A., 2012. Estimated rules for monetary policy. *Federal Reserve Bank of Kansas City Economic Review*, 97(4), p.5-29.

Kazanas, T., Phillippopoulus, A., and Tzavalis, E., 2011. Monetary policy rules and business cycle conditions. *The Manchester School* 79(S), p.73 – 97.

Kim, C., and Nelson, C.R., 2006. Estimation of a forward-looking monetary policy rule: A time-varying parameter model using ex post data. *Journal of Monetary Economics*, 53(8), p.1949-1966.

Kliesen, K.L., Owyang, M.T., and Vermann, E.K., 2012. Disentangling diverse measures: A survey of financial stress indexes. *Federal Reserve Bank of St. Louis Review*, 94(5), p.369-397.

Kontonikas, A., and C., Ioannidis, 2005. Should monetary policy respond to asset price misalignments?. *Economic Modelling*, 22(6), p.1105-1121.

Kontonikas, A., and Kostakis, A., 2013. On monetary policy and stock market anomalies. *Journal of Business Finance and Accounting*, 40(7-8), p.1009-1042.

Kontonikas, A., and Montagnoli, A., 2006. Optimal monetary policy and asset price misalignments. *Scottish Journal of Political Economy*, 53(5), p.636-654.

Kozicki, S., 1999. How useful are Taylor rules for monetary policy?. *Federal Reserve Bank of Kansas City Economic Review*, 84(2), p.5-33.

Kuttner, K.N., and Mosser, P.C., 2002. The monetary transmission mechanism: Some answers and further questions. *Federal Reserve Bank of New York Economic Policy Review*, 8(1), p.15-26.

Lacker, J.M., 2004. Payment system disruptions and the Federal Reserve following September 11, 2001. *Journal of Monetary Economics*, 51(5), p.935-965.

Levin, A., Wieland, V., and Williams, J.C., 2003. The performance of forecast-based monetary policy rules under model uncertainty. *The American Economic Review*, 93(3), p.622-649.

Lombardi, M., and Zhu, F., 2014. A shadow policy rate to calibrate US monetary policy at the zero lower bound. *BIS Working Papers*, Working paper 452.

Martin, C., and Milas, C., 2010. Testing the opportunistic approach to monetary policy. *The Manchester School*, 78(2), p.110-125.

Martin, C., and Milas, C., 2013. Financial crises and monetary policy: Evidence from the UK. *Journal of Financial Stability*, 9(4), p.654-661.

Mattesini, F., and Becchetti, L., 2009. The stock market and the Fed. *Applied Financial Economics*, 19(2), p.99-110.

Mavroeidis, S., 2004. Weak identification of forward-looking models in monetary economics. Oxford Bulletin of Economics and Statistics, 66(S), p.609-635.

McCallum, B.T., 1988. Robustness properties of a rule for monetary policy. *Carnegie-Rochester Conference Series on Public Policy*, 29(1), p.173-203.

Mehra, Y.P., and Minton, B.D., 2007. A Taylor rule and Greenspan era. *Federal Reserve Bank of Richmond Economic Quarterly*, 93(3), p.229-250.

Mehra, Y.P., and Sawhney, B., 2010. Inflation measure, Taylor rules, and the Greenspan-Bernanke years. *Federal Reserve Bank of Richmond Economic Quarterly*, 96(2), p.123-151.

Milas, C., and Naraidoo, R., 2012. Financial conditions and nonlinearities in the European Central Bank (ECB) reaction function: In and out-of-sample assessment. *Computational Statistics and Data Analysis*, 56(1), p.173-189.

Miller, M, Weller, P., and Zhang, L., 2002. Moral hazard and the US stock market: Analysing the ‘Greenspan Put’. *The Economic Journal*, 112(478), p.171-186.

Mishkin, F.S., 2008. How should we respond to asset price bubbles? *Banque de France Financial Stability Review*, 12, p.65-74.

Mishkin, F.S., 2011. Monetary policy strategy: Lessons from the crisis. *NBER Working Working Paper Series*, Working paper 16755.

Mishkin, F.S., and White, E.N., 2002. E.U. stock market crashes and their aftermath: Implications for monetary policy. *NBER Working Paper Series*, Working paper 8992.

Molodtsova, T., Nikolsko-Rzhevskyy, A., and Papell, D.H., 2008. Taylor rules with real-time data: A tale of two countries and one exchange rate. *Journal of Monetary Economics*, 55(S), p.63-79.

Montagnoli, A., and Napolitano, O., 2005. Financial condition index and interest rate settings: A comparative analysis. *University of Naples "Parthenope" Department of Economic Studies*, Working paper 8.2005.

Neely, C.J., 2004. The Federal Reserve responds to crises: September 11th was not the first. *Federal Reserve Bank of St. Louis Review*, 86(2), p.27-42.

Nikolsko-Rzhevskyy, A., 2011. Monetary policy estimations in real time: Forward-looking Taylor rules without forward-looking data. *Journal of Money, Credit and Banking*, 43(5), p.873-897.

Nobay, A.R., and Peel, D.A., 2003. Optimal discretionary monetary policy in a model of asymmetric central bank preferences. *The Economic Journal*, 113(489), p.657-665.

Orphanides, A., 2001. Monetary policy rules based on real-time data. *The American Economic Review*, 91(4), p.964-985.

Orphanides, A., 2007. Taylor rules. *Federal Reserve Board Finance and Economics Discussions Series*, Working paper 18.

Orphanides, A., and Wieland, V., 2008. Economic projections and rules of thumb for monetary policy. *Federal Reserve Bank of St. Louis Review*, 90(4), p.307-324.

Osterholm, P., 2005. The Taylor rule and real-time data – a critical appraisal. *Applied Economics Letters*, 12(11), p.679-685.

Pastor, L., and Stambaugh, R.F., 2003. Liquidity risk and expected stock returns. *Journal of Political Economy*, 111(3), p.642-685.

Pavasuthipaisit, R., 2010. The role of asset prices in best-practice monetary policy. *International Journal of Central Banking*, 6(2), p.81-115.

Ravn, S.H., 2012. Has the Fed reacted asymmetrically to stock prices? *The B.E Journal of Macroeconomics*, 12(1), Article 14, p.1-34.

Rigobon, R., and Sack, B., 2003. Measuring the reaction of monetary policy to the stock market. *The Quarterly Journal of Economics*, 118(2), p.639-669.

Roubini, N., 2006. Why central banks should burst bubbles. *International Finance*, 9(1), p.87-107.

Rudebusch, G.D., 2002. Term structure evidence on interest rate smoothing and monetary policy inertia. *Journal of Monetary Economics*, 49(6), p.1161-1187.

Sack, B., 2000. Does the Fed act gradually? A VAR analysis. *Journal of Monetary Economics*, 46(1), p.229-256.

Sauer, S., 2007. Three liquidity crises in retrospective: Implications for central banking today. *University of Munich Department of Economics*, Discussion paper 2007-26.

Schaling, E., 2004. The nonlinear Phillips curve and inflation forecast targeting: Symmetric versus asymmetric monetary policy rules. *Journal of Money, Credit and Banking*, 36(3), p.361-386.

Schularick, M., and Taylor, A.M., 2012. Credit booms gone bust: Monetary policy, leverage cycles and financial crises. *The American Economic Review*, 102(2), p.1029-1061.

Seyfried, W.L., and Bremmer, D.S., 2001. Analyzing Fed behavior using a dynamic Taylor-type rule. *Journal of Economics and Finance*, 25(1), p.23-32.

Shea, J., 1997. Instrument relevance in multivariate linear models: A simple measure. *The Review of Economics and Statistics*, 79(2), p.348-352.

Siklos, P.L., Werner, T., and Bohl, M.T., Asset prices in Taylor rules: Specification, estimation, and policy implications for the ECB. *Deutsche Bundesbank*, Discussion paper 22/2004.

Sousa, R.M., 2014. Wealth, asset portfolio, money demand and policy rule. *Bulletin of Economic Research*, 66(1), p.95-111.

Staiger, D., and Stock, J.H., 1997. Instrumental variables regression with weak instruments. *Econometrica*, 65(3), p.557-586.

Stock, J.H., and Yogo, M., 2005. Testing for weak instruments in linear IV regression. In *Identification and Inference for Econometric Models: Essays in Honor of Thomas Rothenberg*, p.80–108, Cambridge University Press.

Surico, P., 2007. The Fed's monetary policy rule and U.S. inflation: The case of asymmetric preferences. *Journal of Economic Dynamics and Control*, 31(1), p.305-324.

Svensson, L.E.O., 1997. Inflation forecast targeting: Implementing and monitoring inflation targets. *European Economic Review*, 41(6), p.1111-1146.

Svensson, L.E.O., 2016. Cost-benefit analysis of leaning against the wind: Are costs larger also with less effective macroprudential policy?. *NBER Working Paper Series*, Working paper 21902.

Taylor, J.B., 1993. Discretion versus policy rules in practice. *Carnegie-Rochester Conference Series on Public Policy*, 39(1), p.195-214.

Taylor, J.B., 1999. A historical analysis of monetary policy rules. In John B. Taylor, 1999. *Monetary Policy Rules*. London: University of Chicago Press. Ch. 7.

Taylor, J.B., 2008. Monetary policy and the state of the economy. Testimony before the Committee on Financial Services U.S. House of Representatives, 26 February, 2008.

Taylor, J.B., 2012. Monetary policy rules work and discretion doesn‘t: A tale of two eras. *Journal of Money, Credit and Banking*, 44(6), p.1017-1032.

Taylor, J.B., and Williams, J.C., 2010. Simple and robust rules for monetary policy. *NBER Working Paper Series*, Working paper 15908.

Teranishi, Y., 2012. Credit spread and monetary policy. *Economics Letters*, 114(1), p.26-28.

Valente, G., 2003. Monetary policy rules and regime shifts. *Applied Financial Economics*, 13(7), p.525-535.

Vasicek, B., 2012. Is monetary policy in the new EU member states asymmetric?. *Economic Systems*, 36(2), p.235-263.

Vickers, J., 2000. Monetary policy and asset prices. *The Manchester School*, 68(S), p.1-22.

Wadhwani, S., 2008. Should monetary policy respond to asset price bubbles? Revisiting the debate. *National Institute Economic Review*, 206(1), p.25-34.

White, W.R., 2004. Are changes in financial structure extending safety nets?. *BIS Working Papers*, Working paper 145.

Wu, J.C., and Xia, F.D., 2016. Measuring the macroeconomic impact of monetary policy at the zero lower bound. *Journal of Money, Credit and Banking*, 48(2-3), p.253-291.

Chapter 1 – Appendix A

Table A1.1: Basic and augmented Taylor rules – direct reaction to financial markets – GMM estimation

	Panel A:	**Panel B: Augmented Taylor rules**			
	Basic Taylor rule	*CFCI*	*CSPR*	*SP*	*LQ*
α	0.274 (*1.348*)	-0.747 (*1.117*)	2.576* (*1.335*)	-0.199 (*0.980*)	2.020 (*1.972*)
ρ_1	1.455*** (*0.010*)	1.455*** (*0.010*)	1.361*** (*0.103*)	1.400*** (*0.102*)	1.498*** (*0.114*)
ρ_2	-0.551*** (*0.085*)	-0.568*** (*0.085*)	-0.477*** (*0.089*)	-0.524*** (*0.084*)	-0.595*** (*0.092*)
β	1.683*** (*0.501*)	2.004*** (*0.427*)	2.136*** (*0.406*)	1.601*** (*0.366*)	1.476** (*0.561*)
γ	1.222** (*0.533*)	1.076** (*0.435*)	0.936** (*0.429*)	1.010*** (*0.381*)	0.795* (*0.467*)
μ	-	1.094** (*0.465*)	-3.92*** (*0.9525*)	0.055** (*0.023*)	0.514* (*0.305*)
Eq. SE	0.365	0.337	0.332	0.351	0.377
$\overline{R}^2$	0.973	0.977	0.978	0.975	0.972
J-statistic	0.207	0.247	0.273	0.344	0.420
KP *F*-statistic	25.739	12.998	20.843	24.058	2.151
PR^2 $(\boldsymbol{\pi_{t+4}})$	0.833	0.808	0.787	0.798	0.838
PR^2 $(\boldsymbol{\hat{y}_t})$	0.738	0.712	0.733	0.718	0.660
PR^2 $(\boldsymbol{x_t})$	-	0.771	0.733	0.749	0.230

Notes: This table reports the GMM estimates of Equation (1.3) in Panel A and Equation (1.4) in Panel B over the sample period 1985:Q1 – 2008:Q4. The financial indicators included into the augmented Taylor rule one by one are: the Citi Financial Conditions Index (*CFCI*), the credit spread between Moody's BAA and AAA corporate bonds (*CRSP*), annual stock returns on the S&P500 index (*SP*), and the stock market liquidity measure by Pastor and Stambaugh (2003) (*LQ*). Estimates are obtained using HAC (Newey-West) weighting matrix and a Bartlett kernel with Newey-West bandwidth selection. The list of instruments includes a constant and four lags of expected inflation, output gap, interest rate, 10-year US Treasury yield and an additional variable when applicable. White heteroscedasticity-consistent standard errors are reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\overline{R}^2$ denotes adjusted R-squared. The values reported for *J*-statistic are the p-values. KP *F*-statistic denotes the statistic for Kleibergen-Paap robust rk Wald F test. PR^2 presents Shea's (1997) partial R^2 measure for each endogenous variable.

Table A1.2: Augmented Taylor rules – direct reaction to financial markets across the business cycle – GMM estimation

	CFCI	*CSPR*	*SP*	*LQ*
α	-0.257 (*0.889*)	1.128 (*1.120*)	1.798 (*1.520*)	1.463 (*1.120*)
ρ_1	1.299*** (*0.087*)	1.320*** (*0.079*)	1.321*** (*0.119*)	1.258*** (*0.117*)
ρ_2	-0.426*** (*0.070*)	-0.433*** (*0.067*)	-0.429*** (*0.100*)	-0.405*** (*0.099*)
β	2.024*** (*0.333*)	2.048*** (*0.386*)	1.478*** (*0.432*)	1.400*** (*0.364*)
γ	1.243*** (*0.350*)	1.138*** (*0.381*)	1.383** (*0.592*)	1.777*** (*0.389*)
μ^{ND}	-0.372 (*0.445*)	-1.685 (*1.077*)	-0.043 (*0.052*)	0.038 (*0.135*)
μ^{D}	3.720** (*1.436*)	-5.312*** (*1.538*)	0.500* (*0.260*)	1.053** (*0.426*)
Eq. SE	0.332	0.288	0.367	0.401
$\bar{R}^2$	0.978	0.983	0.973	0.968
$\boldsymbol{WT^X}$	0.001	0.000	0.024	0.003
J-statistic	0.300	0.241	0.271	0.531
KP *F*-statistic	0.889	2.325	1.162	1.846
PR^2 ($\boldsymbol{\pi_{t+4}}$)	0.815	0.803	0.785	0.799
PR^2 ($\boldsymbol{\hat{y}_t}$)	0.677	0.665	0.706	0.447
PR^2 ((1-D^R)$\boldsymbol{x_t}$)	0.506	0.718	0.511	0.215
PR^2 ($D^R\boldsymbol{x_t}$)	0.386	0.711	0.298	0.247

Notes: This table reports the GMM estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator in economic expansions and economic recessions, respectively. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The penultimate row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^{D}$. See also Table A1.1 notes.

Table A1.3: Basic Taylor rules - indirect reaction to financial markets – GMM estimation

	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}
α	0.017 (*1.372*)	-0.159 (*1.702*)	-0.841 (*1.722*)	1.444 (*1.906*)
ρ_1	1.445*** (*0.112*)	1.481*** (*0.144*)	1.460*** (*0.127*)	1.482*** (*0.130*)
ρ_2	-0.547*** (*0.100*)	-0.572*** (*0.123*)	-0.566*** (*0.113*)	-0.595*** (*0.112*)
β^{ND}	1.925*** (*0.597*)	2.032*** (*0.677*)	2.452*** (*0.711*)	1.518** (*0.669*)
β^{D}	1.163** (*0.571*)	0.956 (*0.734*)	0.689 (*1.059*)	-1.812 (*2.003*)
γ^{ND}	-0.500 (*1.921*)	0.369 (*0.926*)	0.689 (*1.152*)	-0.077 (*1.113*)
γ^{D}	2.582** (*1.204*)	4.409 (*3.291*)	1.496 (*2.788*)	5.429* (*2.868*)
Eq. SE	0.395	0.388	0.390	0.480
$\bar{R}^2$	0.969	0.970	0.969	0.954
$\boldsymbol{WT^{\pi}}$	0.239	0.236	0.026	0.006
$\boldsymbol{WT^{\hat{y}}}$	0.301	0.243	0.824	0.032
J-statistic	0.159	0.208	0.136	0.169
KP *F*-statistic	0.961	1.004	0.489	0.576
PR^2 ((1-D^X)$\boldsymbol{\pi_{t+4}}$)	0.717	0.749	0.613	0.813
PR^2 ($D^X\boldsymbol{\pi_{t+4}}$)	0.836	0.718	0.617	0.357
PR^2 ((1-D^X)$\boldsymbol{\hat{y}_t}$)	0.207	0.407	0.160	0.253
PR^2 ($D^X\boldsymbol{\hat{y}_t}$)	0.352	0.196	0.075	0.089

Notes: This table reports the GMM estimates of Equation (1.6) over the sample period 1985:Q1 – 2008:Q4 using four financial stress dummy variables (D^X). D^{CFCI} indicates financial stress related to overall financial conditions, D^{CSPR} denotes credit risk-related stress, D^{SP} identifies stock market bear conditions and D^{LQ} denotes stock market liquidity-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.1 notes.

Table A1.4: Basic and augmented Taylor rules – financial crisis effect – GMM estimation

	Panel A: Basic Taylor rule	Panel B: Augmented Taylor rules			
	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*
α	1.021 (*0.856*)	0.201 (*0.886*)	0.778 (*0.962*)	0.264 (*0.948*)	0.932 (*0.899*)
ρ_1	1.043*** (*0.133*)	1.223*** (*0.133*)	1.270*** (*0.139*)	1.323*** (*0.125*)	1.225*** (*0.134*)
ρ_2	-0.240** (*0.105*)	-0.399*** (*0.104*)	-0.418*** (*0.112*)	-0.470*** (*0.097*)	-0.397*** (*0.104*)
β	-	1.741*** (*0.310*)	1.539*** (*0.422*)	1.701*** (*0.340*)	1.561*** (*0.281*)
γ	-	1.225*** (*0.303*)	1.370*** (*0.401*)	1.138*** (*0.380*)	1.575*** (*0.308*)
μ^{ND}	-	-0.083 (*0.303*)	0.020 (*1.179*)	0.008 (*0.022*)	0.033 (*0.118*)
μ^{D}	-	4.058** (*1.614*)	-6.531*** (*2.351*)	0.410** (*0.202*)	0.923*** (*0.340*)
β^{ND}	1.450*** (*0.301*)	-	-	-	-
β^{D}	-3.250* (*1.708*)	-	-	-	-
γ^{ND}	1.744*** (*0.344*)	-	-	-	-
γ^{D}	1.298 (*1.766*)	-	-	-	-
Eq. SE	0.396	0.366	0.367	0.367	0.375
$\overline{R}^2$	0.969	0.973	0.973	0.973	0.972
$\boldsymbol{WT^X}$	-	0.010	0.003	0.029	0.002
$\boldsymbol{WT^\pi}$	0.008	-	-	-	-
$\boldsymbol{WT^{\hat{y}}}$	0.797	-	-	-	-
J-statistic	0.358	0.313	0.335	0.352	0.492
KP *F*-statistic	0.252	0.807	0.661	0.361	1.726
PR^2 ($\boldsymbol{\pi_{t+4}}$)	-	0.780	0.800	0.775	0.760
PR^2 ($\boldsymbol{\hat{y}_t}$)	-	0.540	0.496	0.651	0.498
PR^2 ((1-D^X)$\boldsymbol{x_t}$)	-	0.581	0.587	0.659	0.189
PR^2 ($D^X\boldsymbol{x_t}$)	-	0.323	0.579	0.305	0.317
PR^2 ((1-D^X)$\boldsymbol{\pi_{t+4}}$)	0.750	-	-	-	-
PR^2 ($D^X\boldsymbol{\pi_{t+4}}$)	0.482	-	-	-	-
PR^2 ((1-D^X)$\boldsymbol{\hat{y}_t}$)	0.243	-	-	-	-
PR^2 ($D^X\boldsymbol{\hat{y}_t}$)	0.143	-	-	-	-

Notes: This table reports the GMM estimates of Equation (1.6) in Panel A and Equation (1.5) in Panel B over the sample period 1985:Q1 – 2008:Q4 using the financial crisis dummy variable (D^{07-08}). D^{07-08} takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator before the crisis and during the crisis, respectively. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation before the crisis and during the crisis, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote policy response to output gap before the crisis and during the crisis, respectively. The three rows (WT^X , WT^π and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^D$, $\beta^{ND} = \beta^D$ and $\gamma^{ND} = \gamma^D$, respectively. See also Table A1.1 notes.

Table A1.5: Augmented Taylor rules – direct reaction to financial markets – alternative financial indicators

	NFCI	*CSPR2*	*MAFCI*	*IBSPR*[1]
α	-3.562**	6.118***	-1.753	-0.533
	(1.641)	***(1.814)***	***(1.911)***	***(1.520)***
ρ_1	1.444***	1.306***	1.433***	1.532***
	(0.078)	***(0.080)***	***(0.102)***	***(0.087)***
ρ_2	-0.549***	-0.435***	-0.524***	-0.632***
	(0.064)	***(0.066)***	***(0.088)***	***(0.072)***
β	2.532***	1.485***	2.121***	2.54***
	(0.512)	***(0.415)***	***(0.587)***	***(0.684)***
γ	1.356***	0.801*	0.974	1.326**
	(0.461)	***(0.441)***	***(0.607)***	***(0.540)***
μ	-2.722**	-2.650***	0.585	-2.902
	(1.141)	***(0.647)***	***(0.418)***	***(2.071)***
Eq. SE	0.332	0.306	0.351	0.345
$\bar{R}^2$	0.978	0.981	0.975	0.975

Notes: This table reports the OLS estimates of Equation (1.4) over the sample period 1985:Q1 – 2008:Q4. Standard errors are reported in parentheses. The financial indicators included into the augmented Taylor rule one by one are: the Chicago National Financial Conditions Index (*NFCI*), the credit spread between Moody's BAA corporate and 10-year US Treasury bonds (*CRSP2*), the Macroeconomic Advisers Monetary and Financial Conditions Index (*MAFCI*), and the interest rate spread between 3-month LIBOR and 3-month US Treasury bill (*IBSPR*). Appropriate standard errors are used based on the White heteroscedasticity test and Ljung-Box Q-statistics. White heteroscedasticity-consistent standard errors are reported in *italic*, while heteroscedasticity and autocorrelation-consistent (HAC) standard errors are reported in ***bold italic***. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\bar{R}^2$ denotes adjusted R-squared.
[1]*IBSPR* data starts in 1986:Q1.

Table A1.6: Augmented Taylor rules – direct reaction to financial markets across the business cycle – alternative financial indicators

	NFCI	*CSPR2*	*MAFCI*	*IBSPR*
α	-2.428 *(1.726)*	3.415** (*1.330*)	-0.786 ***(1.850)***	-0.126 *(0.832)*
ρ_1	1.436*** *(0.087)*	1.267*** *(0.067)*	1.417*** ***(0.102)***	1.337*** *(0.082)*
ρ_2	-0.551*** *(0.076)*	-0.390*** *(0.054)*	-0.505*** ***(0.088)***	-0.475*** *(0.069)*
β	2.307*** *(0.478)*	1.553*** *(0.323)*	1.955*** ***(0.597)***	1.593*** *(0.387)*
γ	1.187*** *(0.358)*	1.025*** *(0.258)*	0.966 ***(0.645)***	0.937*** *(0.231)*
μ^{ND}	-1.698 *(1.205)*	-1.269*** *(0.420)*	0.226 ***(0.295)***	1.224 *(0.832)*
μ^{D}	-3.487*** *(1.272)*	-2.736*** *(0.593)*	1.962** ***(0.980)***	-3.376*** *(1.049)*
Eq. SE	0.331	0.272	0.340	0.292
$\overline{R}^2$	0.978	0.985	0.977	0.982
$\boldsymbol{WT^X}$	0.277	0.005	0.046	0.000

Notes: This table reports the OLS estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4. The estimates of μ^{ND} and μ^{D} denote the policy response to financial indicator in economic expansions and economic recessions, respectively. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The last row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^{D}$. See also Table A1.5 notes.

Table A1.7: Basic Taylor rules - indirect reaction to financial markets – alternative financial indicators

	D^{NFCI}	D^{CSPR2}	D^{MAFCI}	D^{IBSPR}
α	-1.934 ***(1.634)***	-0.882 ***(1.264)***	-0.954 ***(1.447)***	-1.677 *(2.164)*
ρ_1	1.457*** ***(0.087)***	1.421*** ***(0.103)***	1.473*** ***(0.119)***	1.535*** *(0.098)*
ρ_2	-0.579*** ***(0.075)***	-0.549*** ***(0.081)***	-0.579*** ***(0.100)***	-0.638*** *(0.086)*
β^{ND}	2.484*** ***(0.623)***	2.334*** ***(0.467)***	2.147*** ***(0.514)***	2.443*** *(0.898)*
β^{D}	2.123*** ***(0.453)***	1.700*** ***(0.420)***	1.740*** ***(0.513)***	2.046*** *(0.527)*
γ^{ND}	1.214*** ***(0.312)***	0.325 ***(0.354)***	0.562 ***(0.677)***	1.115*** *(0.424)*
γ^{D}	0.292 *(1.428)*	1.235** ***(0.586)***	1.768** ***(0.751)***	1.002 *(1.046)*
Eq. SE	0.352	0.344	0.354	0.360
$\bar{R}^2$	0.975	0.976	0.975	0.973
$\boldsymbol{WT^{\pi}}$	0.525	0.007	0.240	0.481
$\boldsymbol{WT^{\hat{y}}}$	0.529	0.095	0.233	0.889

Notes: This table reports the OLS estimates of Equation (1.6) over the sample period 1985:Q1 – 2008:Q4 using four financial stress dummy variables (D^X). D^{NFCI} and D^{MAFCI} indicate financial stress related to overall financial conditions, D^{CSPR2} and D^{IBSPR} denote credit risk-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The last two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.5 notes.

Table A1.8: Basic and augmented Taylor rules – financial crisis effect – alternative financial indicators

	NFCI	*CSPR2*	*MAFCI*	*IBSPR*
α	-1.061 *(1.458)*	3.730** ***(1.600)***	-1.340 ***(1.657)***	0.293 *(0.740)*
ρ_1	1.413*** *(0.086)*	1.273*** ***(0.063)***	1.416*** ***(0.106)***	1.330*** *(0.090)*
ρ_2	-0.536*** *(0.076)*	-0.418*** ***(0.052)***	-0.517*** ***(0.088)***	-0.494*** *(0.076)*
β	2.020*** *(0.420)*	1.532*** ***(0.345)***	2.052*** ***(0.533)***	1.324*** *(0.350)*
γ	1.022*** *(0.315)*	0.989*** ***(0.300)***	0.860 ***(0.535)***	0.937*** *(0.205)*
μ^{ND}	-0.623 *(0.926)*	-1.510*** ***(0.505)***	0.440 ***(0.303)***	1.403* *(0.814)*
μ^{D}	-4.452*** *(1.278)*	-2.549*** ***(0.540)***	3.443 ***(2.310)***	-2.848*** (0.833)
Eq. SE	0.324	0.293	0.346	0.309
$\bar{R}^2$	0.979	0.983	0.976	0.980
$\boldsymbol{WT^X}$	0.011	0.023	0.202	0.000

Notes: This table reports the OLS estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4 using the financial crisis dummy variable ($D^{07\text{-}08}$). $D^{07\text{-}08}$ takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator before the crisis and during the crisis, respectively. The last row (WT^X) represents the Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^{D}$. See also Table A1.5 notes.

Table A1.9: Basic and augmented Taylor rules – direct reaction to financial markets – real-time output gap measure

	Panel A:	Panel B: Augmented Taylor rules			
	Basic Taylor rule	*CFCI*	*CSPR*	*SP*	*LQ*
α	-2.468 *(1.520)*	-2.485* *(1.465)*	0.458 *(1.483)*	-2.319* *(1.296)*	-1.551 *(1.329)*
ρ_1	1.404*** *(0.071)*	1.416*** *(0.075)*	1.336*** *(0.074)*	1.390*** *(0.073)*	1.413*** *(0.072)*
ρ_2	-0.495*** *(0.064)*	-0.515*** *(0.066)*	-0.439*** *(0.065)*	-0.492*** *(0.065)*	-0.505*** *(0.064)*
β	2.546*** *(0.523)*	2.583*** *(0.523)*	2.714*** *(0.456)*	2.410*** *(0.446)*	2.368*** *(0.483)*
γ	1.685*** *(0.553)*	1.416*** *(0.503)*	1.288*** *(0.406)*	1.385*** *(0.475)*	1.594*** *(0.520)*
μ	-	0.640 (*0.603*)	-3.535** *(1.606)*	0.028 *(0.034)*	0.157 *(0.129)*
Eq. SE	0.338	0.335	0.322	0.338	0.334
$\bar{R}^2$	0.977	0.978	0.979	0.977	0.978

Notes: This table reports the OLS estimates of Equation (1.3) in Panel A and Equation (1.4) in Panel B over the sample period 1985:Q1 – 2008:Q4. The financial indicators included into the augmented Taylor rule one by one are: the Citi Financial Conditions Index (*CFCI*), the credit spread between Moody's BAA and AAA corporate bonds (*CRSP*), annual stock returns on the S&P500 index (*SP*), and the stock market liquidity measure by Pastor and Stambaugh (2003) (*LQ*). Appropriate standard errors are used based on the White heteroscedasticity test and Ljung-Box Q-statistics. White heteroscedasticity-consistent standard errors are reported in *italic*, while heteroscedasticity and autocorrelation-consistent (HAC) standard errors are reported in ***bold italic***. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\bar{R}^2$ denotes adjusted R-squared.

Table A1.10: Augmented Taylor rules – direct reaction to financial markets across the business cycle – real-time output gap measure

	CFCI	*CSPR*	*SP*	*LQ*
α	-1.703 *(1.307)*	-0.875 *(1.412)*	-1.048 *(1.343)*	-1.397 *(1.358)*
ρ_1	1.346*** *(0.075)*	1.339*** *(0.069)*	1.355*** *(0.071)*	1.368*** *(0.074)*
ρ_2	-0.437*** *(0.066)*	-0.432*** *(0.058)*	-0.440*** *(0.064)*	-0.462*** *(0.066)*
β	2.462*** *(0.499)*	2.374*** *(0.454)*	2.283*** *(0.486)*	2.267*** *(0.483)*
γ	1.441*** *(0.462)*	1.136*** *(0.402)*	1.657*** *(0.567)*	1.579*** *(0.468)*
μ^{ND}	-0.548 *(0.390)*	-0.525 *(1.186)*	-0.034 *(0.037)*	0.000 *(0.071)*
μ^{D}	2.809* *(1.439)*	-4.629*** *(1.450)*	0.302** *(0.147)*	0.528** *(0.265)*
Eq. SE	0.307	0.291	0.313	0.315
$\bar{R}^2$	0.981	0.983	0.980	0.980
$\boldsymbol{WT^X}$	0.035	0.017	0.052	0.046

Notes: This table reports the OLS estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator in economic expansions and economic recessions, respectively. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The last row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^{D}$. See also Table A1.9 notes.

Table A1.11: Basic Taylor rules - indirect reaction to financial markets – real-time output gap measure

	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}
α	-2.338 *(1.552)*	-3.330** *(1.690)*	-2.218 *(1.490)*	-1.588 *(1.240)*
ρ_1	1.359*** *(0.083)*	1.345*** *(0.076)*	1.398*** *(0.075)*	1.397*** *(0.074)*
ρ_2	-0.451*** *(0.076)*	-0.435*** *(0.070)*	-0.492*** *(0.067)*	-0.497*** *(0.065)*
β^{ND}	2.484*** *(0.583)*	2.991*** *(0.624)*	2.517*** *(0.543)*	2.317*** *(0.449)*
β^{D}	2.596*** *(0.518)*	2.090*** *(0.502)*	2.316*** *(0.522)*	1.902** *(0.760)*
γ^{ND}	1.308** *(0.558)*	1.921*** *(0.617)*	1.532*** *(0.551)*	1.386*** *(0.460)*
γ^{D}	2.374*** *(0.899)*	1.291 *(1.006)*	1.782 (1.388)	2.819 *(1.768)*
Eq. SE	0.339	0.330	0.341	0.337
$\overline{R}^2$	0.977	0.978	0.977	0.977
$\boldsymbol{WT^{\pi}}$	0.705	0.057	0.681	0.523
$\boldsymbol{WT^{\hat{y}}}$	0.257	0.575	0.864	0.425

Notes: This table reports the OLS estimates of Equation (6) over the sample period 1985:Q1 – 2008:Q4 using four financial stress dummy variables (D^X). D^{CFCI} indicates financial stress related to overall financial conditions, D^{CSPR} denotes credit risk-related stress, D^{SP} identifies stock market bear conditions and D^{LQ} denotes stock market liquidity-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The last two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.9 notes.

Table A1.12: Basic and augmented Taylor rules – financial crisis effect – real-time output gap measure

	Panel A: Basic Taylor rule	Panel B: Augmented Taylor rules			
	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*
α	-1.412 *(1.107)*	-1.352 *(1.103)*	-0.797 *(1.337)*	-1.618 *(1.112)*	-1.317 *(1.204)*
ρ_1	1.327*** *(0.075)*	1.335*** *(0.072)*	1.331*** *(0.072)*	1.364*** *(0.073)*	1.360*** *(0.074)*
ρ_2	-0.432*** *(0.066)*	-0.441*** *(0.065)*	-0.437*** *(0.061)*	-0.470*** *(0.064)*	-0.460*** *(0.064)*
β	-	2.242*** *(0.411)*	2.297*** *(0.425)*	2.315*** *(0.409)*	2.244*** *(0.436)*
γ	-	1.448*** *(0.421)*	1.34*** *(0.399)*	1.398*** *(0.453)*	1.495*** *(0.425)*
μ^{ND}	-	-0.108 *(0.318)*	-0.732 *(1.275)*	-0.000 *(0.025)*	0.004 *(0.071)*
μ^{D}	-	2.603* *(1.394)*	-3.803*** *(1.195)*	0.259** *(0.103)*	0.572** *(0.243)*
β^{ND}	2.267*** *(0.413)*	-	-	-	-
β^{D}	0.796 *(1.226)*	-	-	-	-
γ^{ND}	1.452*** *(0.392)*	-	-	-	-
γ^{D}	5.923 *(5.886)*	-	-	-	-
Eq. SE	0.318	0.317	0.309	0.320	0.311
$\bar{R}^2$	0.980	0.980	0.981	0.980	0.981
$\boldsymbol{WT^X}$	-	0.064	0.051	0.023	0.022
$\boldsymbol{WT^\pi}$	0.195	-	-	-	-
$\boldsymbol{WT^{\hat{y}}}$	0.448	-	-	-	-

Notes: This table reports the OLS estimates of Equation (1.6) in Panel A and Equation (1.5) in Panel B over the sample period 1985:Q1 – 2008:Q4 using the financial crisis dummy variable ($D^{07\text{-}08}$). $D^{07\text{-}08}$ takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator before the crisis and during the crisis, respectively. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation before the crisis and during the crisis, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote the policy response to output gap before the crisis and during the crisis, respectively. The last three rows (WT^X , WT^π and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^{D}$, $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.9 notes.

Table A1.13: Basic and augmented Taylor rules – direct reaction to financial markets – alternative output gap measure (quadratic trend)

	Panel A:	Panel B: Augmented Taylor rules			
	Basic Taylor rule	*CFCI*	*CSPR*	*SP*	*LQ*
α	-4.347**	-4.977***	-0.990	-4.039***	-3.334*
	(2.168)	*(1.562)*	***(1.552)***	*(1.337)*	*(1.687)*
ρ_1	1.473***	1.427***	1.337***	1.367***	1.465***
	(0.083)	*(0.084)*	***(0.078)***	*(0.095)*	*(0.086)*
ρ_2	-0.570***	-0.545***	-0.457***	-0.507***	-0.567***
	(0.082)	*(0.077)*	***(0.079)***	*(0.084)*	*(0.082)*
β	3.124***	3.358***	3.324***	2.821***	2.947***
	(0.687)	*(0.517)*	***(0.530)***	*(0.433)*	*(0.587)*
γ	0.484*	0.595***	0.476**	0.461***	0.527**
	(0.267)	*(0.181)*	***(0.195)***	*(0.141)*	*(0.225)*
μ	-	1.233***	-4.114***	0.062***	0.211*
		(0.401)	***(1.050)***	*(0.020)*	*(0.116)*
Eq. SE	0.360	0.332	0.328	0.342	0.349
$\bar{R}^2$	0.974	0.978	0.978	0.976	0.976

Notes: This table reports the OLS estimates of Equation (1.3) in Panel A and Equation (1.4) in Panel B over the sample period 1985:Q1 – 2008:Q4. The financial indicators included into the augmented Taylor rule one by one are: the Citi Financial Conditions Index (*CFCI*), the credit spread between Moody's BAA and AAA corporate bonds (*CRSP*), annual stock returns on the S&P500 index (*SP*), and the stock market liquidity measure by Pastor and Stambaugh (2003) (*LQ*). Appropriate standard errors are used based on the White heteroscedasticity test and Ljung-Box Q-statistics. White heteroscedasticity-consistent standard errors are reported in *italic*, while heteroscedasticity and autocorrelation-consistent (HAC) standard errors are reported in ***bold italic***. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\bar{R}^2$ denotes adjusted R-squared.

Table A1.14: Augmented Taylor rules – direct reaction to financial markets across the business cycle – alternative output gap measure (quadratic trend)

	CFCI	*CSPR*	*SP*	*LQ*
α	-3.939***	-2.091	-3.147**	-2.930*
	(1.466)	***(1.575)***	*(1.345)*	*(1.592)*
ρ_1	1.400***	1.337***	1.360***	1.438***
	(0.076)	***(0.060)***	*(0.080)*	*(0.088)*
ρ_2	-0.505***	-0.444***	-0.478***	-0.538***
	(0.069)	***(0.053)***	*(0.070)*	*(0.084)*
β	3.129***	2.960***	2.701***	2.783***
	(0.511)	***(0.559)***	*(0.444)*	*(0.552)*
γ	0.481***	0.433**	0.446***	0.469**
	(0.156)	***(0.195)***	*(0.141)*	*(0.185)*
μ^{ND}	0.457	-1.304	0.027	0.083
	(0.421)	***(1.189)***	(*0.026*)	*(0.079)*
μ^{D}	2.693**	-5.033***	0.234**	0.522**
	(1.153)	***(1.637)***	(*0.102*)	*(0.243)*
Eq. SE	0.318	0.295	0.326	0.335
$\bar{R}^2$	0.980	0.983	0.979	0.977
$\boldsymbol{WT^X}$	0.065	0.005	0.082	0.059

Notes: This table reports the OLS estimates of Equation (1.5) over the sample period 1985:Q1 – 2008:Q4. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator in economic expansions and economic recessions, respectively. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The last row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^{D}$. See also Table A1.13 notes.

Table A1.15: Basic Taylor rules - indirect reaction to financial markets - alternative output gap measure (quadratic trend)

	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}
α	-5.972**	-5.321*	-4.308**	-3.259**
	(2.413)	***(2.765)***	*(2.019)*	*(1.584)*
ρ_1	1.483***	1.463***	1.400***	1.453***
	(0.082)	***(0.084)***	*(0.080)*	*(0.089)*
ρ_2	-0.581***	-0.549***	-0.512***	-0.558***
	(0.078)	***(0.084)***	*(0.075)*	*(0.083)*
β^{ND}	3.997***	3.618***	3.244***	2.856***
	(0.890)	***(0.980)***	*(0.700)*	*(0.548)*
β^{D}	3.261***	2.804***	2.609***	1.142
	(0.654)	***(0.864)***	*(0.558)*	*(1.245)*
γ^{ND}	0.685***	0.442	0.602***	0.411**
	(0.248)	***(0.286)***	*(0.176)*	*(0.169)*
γ^{D}	0.672*	1.268	0.278	1.177
	(0.375)	***(0.995)***	*(0.630)*	*(0.773)*
Eq. SE	0.351	0.351	0.349	0.350
$\bar{R}^2$	0.975	0.975	0.976	0.976
$\boldsymbol{WT^{\pi}}$	0.073	0.131	0.232	0.132
$\boldsymbol{WT^{\hat{y}}}$	0.967	0.370	0.574	0.317

Notes: This table reports the OLS estimates of Equation (1.6) over the sample period 1985:Q1 – 2008:Q4 using four financial stress dummy variables (D^X). D^{CFCI} indicates financial stress related to overall financial conditions, D^{CSPR} denotes credit risk-related stress, D^{SP} identifies stock market bear conditions and D^{LQ} denotes stock market liquidity-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The last two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.13 notes.

Table A1.16: Basic and augmented Taylor rules – financial crisis effect - alternative output gap measure (quadratic trend)

	Panel A: Basic Taylor rule	**Panel B: Augmented Taylor rules**			
	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*
α	-2.887* ***(1.575)***	-4.118*** *(1.368)*	-1.293 ***(1.535)***	-3.332*** *(1.252)*	-2.716* *(1.473)*
ρ_1	1.406*** ***(0.092)***	1.410*** *(0.079)*	1.351*** ***(0.080)***	1.377*** *(0.088)*	1.435*** *(0.083)*
ρ_2	-0.515*** ***(0.087)***	-0.529*** *(0.074)*	-0.470*** ***(0.080)***	-0.511*** *(0.079)*	-0.538*** *(0.078)*
β	-	3.105*** *(0.461)*	3.050*** ***(0.543)***	2.674*** *(0.421)*	2.719*** *(0.515)*
γ	-	0.514*** (0.157)	0.424** ***(0.202)***	0.388*** *(0.133)*	0.432** *(0.178)*
μ^{ND}	-	0.920*** *(0.288)*	-2.807** ***(1.389)***	0.046** *(0.019)*	0.083 *(0.840)*
μ^{D}	-	2.096* *(1.241)*	-4.369*** ***(1.194)***	0.185** *(0.081)*	0.560** *(0.230)*
β^{ND}	2.722*** ***(0.509)***	-	-	-	-
β^{D}	-0.069 ***(1.123)***	-	-	-	-
γ^{ND}	0.390* ***(0.210)***	-	-	-	-
γ^{D}	1.378** ***(0.647)***	-	-	-	-
Eq. SE	0.339	0.329	0.326	0.336	0.333
$\overline{R}^2$	0.977	0.978	0.979	0.977	0.978
$\boldsymbol{WT^X}$	-	0.357	0.199	0.117	0.039
$\boldsymbol{WT^\pi}$	0.006	-	-	-	-
$\boldsymbol{WT^{\hat{y}}}$	0.139	-	-	-	-

Notes: This table reports the OLS estimates of Equation (1.6) in Panel A and Equation (1.5) in Panel B over the sample period 1985:Q1 – 2008:Q4 using the financial crisis dummy variable ($D^{07\text{-}08}$). $D^{07\text{-}08}$ takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. The estimates of μ^{ND} and μ^{D} denote the policy response to a financial indicator before the crisis and during the crisis, respectively. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation before the crisis and during the crisis, respectively. The estimates of $\gamma^{ND}$and γ^{D} denote thepolicy response to output gap before the crisis and during the crisis, respectively. The last three rows (WT^X , WT^π and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^{D}$, $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.13 notes.

Table A1.17: Basic and augmented Taylor rules – direct reaction to financial markets – alternative sample periods

	Panel A: Basic Taylor rule		Panel B: Augmented Taylor rules							
	1979:Q4 – 2008:Q4	1987:Q4 – 2008:Q4	1979:Q4 – 2008:Q4				1987:Q4 – 2008:Q4			
			CFCI	*CSPR*	*SP*	*LQ*	*CFCI*	*CSPR*	*SP*	*LQ*
α	-0.837	-0.639	-1.238	0.454	-1.194*	-0.684	-1.321	2.848**	-1.019	0.398
	(0.754)	*(1.425)*	*(0.849)*	***(0.849)***	***(0.685)***	***(0.666)***	*(1.260)*	*(1.324)*	*(1.161)*	*(1.301)*
ρ_1	0.716***	1.569***	1.319***	0.732***	0.700***	0.713***	1.511***	1.426***	1.485***	1.558***
	(0.068)	*(0.099)*	*(0.134)*	***(0.067)***	***(0.061)***	***(0.068)***	*(0.101)*	*(0.097)*	*(0.105)*	*(0.099)*
ρ_2	-	-0.672***	-0.471***	-	-	-	-0.632***	-0.559***	-0.619***	-0.665***
		(0.091)	*(0.100)*				*(0.091)*	*(0.088)*	*(0.094)*	*(0.091)*
β	1.986***	1.862***	2.136***	2.341***	1.946***	1.984***	2.194***	1.964***	1.872***	1.668***
	(0.193)	*(0.537)*	*(0.350)*	***(0.284)***	***(0.186)***	***(0.188)***	*(0.481)*	*(0.392)*	*(0.425)*	*(0.499)*
γ	0.961***	0.891*	0.864***	0.767**	0.878***	0.966***	1.037***	0.797***	0.772**	0.966**
	(0.257)	*(0.454)*	*(0.268)*	***(0.313)***	***(0.241)***	***(0.255)***	*(0.374)*	*(0.296)*	*(0.330)*	*(0.406)*
μ	-	-	0.944***	-2.303**	0.051***	0.052	1.094**	-4.085***	0.053**	0.186
			(0.291)	***(1.148)***	***(0.017)***	***(0.066)***	*(0.450)*	*(0.997)*	*(0.025)*	*(0.119)*
Eq. SE	0.853	0.358	0.411	0.830	0.821	0.854	0.335	0.320	0.346	0.348
$\bar{R}^2$	0.947	0.973	0.973	0.950	0.951	0.947	0.977	0.979	0.975	0.975

Notes: This table reports the OLS estimates of Equation (1.3) in Panel A and Equation (1.4) in Panel B over the sample periods 1979:Q4 – 2008:Q4 and 1987:Q4 – 2008:Q4. The financial indicators included into the augmented Taylor rule one by one are: the Citi Financial Conditions Index (*CFCI*), the credit spread between Moody's BAA and AAA corporate bonds (*CRSP*), annual stock returns on the S&P500 index (*SP*), and the stock market liquidity measure by Pastor and Stambaugh (2003) (*LQ*). Appropriate standard errors are used based on the White heteroscedasticity test and Ljung-Box Q-statistics. White heteroscedasticity-consistent standard errors are reported in *italic*, while heteroscedasticity and autocorrelation-consistent (HAC) standard errors are reported in ***bold italic***. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively. Eq. SE denotes the standard error of regression and $\bar{R}^2$ denotes adjusted R-squared.

Table A1.18: Augmented Taylor rules – direct reaction to financial markets across the business cycle – alternative sample periods

	1979:Q4 – 2008:Q4				**1987:Q4 – 2008:Q4**			
	CFCI	*CSPR*	*SP*	*LQ*	*CFCI*	*CSPR*	*SP*	*LQ*
α	-1.041 *(0.904)*	-1.710* ***(1.009)***	-1.470** ***(0.626)***	-1.004 ***(0.671)***	-0.724 *(1.190)*	0.369 *(1.516)*	0.032 *(1.281)*	0.501 *(1.217)*
ρ_1	1.267*** *(0.127)*	0.769*** ***(0.057)***	0.687*** ***(0.071)***	0.721*** ***(0.062)***	1.440*** *(0.092)*	1.372*** *(0.080)*	1.455*** *(0.099)*	1.463*** *(0.108)*
ρ_2	-0.416*** *(0.091)*	-	-	-	-0.551*** *(0.081)*	-0.486*** *(0.067)*	-0.566*** *(0.090)*	-0.582*** *(0.100)*
β	2.169*** *(0.363)*	2.395*** ***(0.259)***	1.975*** ***(0.155)***	2.063*** ***(0.184)***	2.084*** *(0.478)*	1.905*** *(0.430)*	1.687*** *(0.477)*	1.574*** *(0.456)*
γ	0.772*** *(0.218)*	0.725** ***(0.316)***	0.882*** ***(0.234)***	1.016*** ***(0.256)***	1.023*** *(0.325)*	1.051*** *(0.294)*	0.830** *(0.405)*	1.188*** *(0.303)*
μ^{ND}	0.428* *(0.244)*	0.265 ***(1.352)***	0.063*** ***(0.023)***	-0.063 ***(0.053)***	0.247 *(0.264)*	-0.556 *(1.122)*	0.014 *(0.026)*	0.030 *(0.053)*
μ^{D}	2.186** *(0.894)*	-2.885*** ***(1.072)***	-0.005 ***(0.095)***	0.277*** ***(0.101)***	2.678** *(1.182)*	-4.462*** *(1.191)*	0.235* *(0.119)*	0.531** *(0.203)*
Eq. SE	0.395	0.755	0.818	0.827	0.312	0.284	0.328	0.318
$\bar{R}^2$	0.975	0.959	0.951	0.950	0.980	0.983	0.978	0.979
$\boldsymbol{WT^X}$	0.075	0.005	0.526	0.003	0.054	0.014	0.101	0.011

Notes: This table reports the OLS estimates of Equation (1.5) over the sample periods 1979:Q4 – 2008:Q4 and 1987:Q4 – 2008:Q4. D^R takes value one in recessions and zero otherwise. The NBER dates are used to define US recessionary periods. The estimates of μ^{ND} and μ^D denote the policy response to financial indicator in economic expansions and economic recessions, respectively. The last row (WT^X) shows the Wald test p-values of the null hypothesis that $\mu^{ND} = \mu^D$. See also Table A1.17 notes.

Table A1.19: Basic Taylor rules - indirect reaction to financial markets – alternative sample periods

	1979:Q4 – 2008:Q4				**1987:Q4 – 2008:Q4**			
	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}	D^{CFCI}	D^{CSPR}	D^{SP}	D^{LQ}
α	-1.705 *(1.208)*	-1.559* ***(1.014)***	-0.755 ***(0.741)***	-0.878 ***(0.812)***	-1.809 *(1.798)*	-0.529 *(1.196)*	-0.832 *(1.439)*	0.109 *(1.292)*
ρ_1	1.383*** *(0.130)*	0.712*** ***(0.072)***	0.723*** ***(0.068)***	0.720*** ***(0.071)***	1.560*** *(0.100)*	1.454*** (*0.104*)	1.536*** *(0.089)*	1.552*** *(0.102)*
ρ_2	-0.515*** *(0.099)*	-	-	-	-0.660*** *(0.092)*	-0.571*** (*0.093*)	-0.644*** *(0.083)*	-0.660*** *(0.092)*
β^{ND}	2.494*** *(0.461)*	2.315*** ***(0.332)***	2.014*** ***(0.234)***	1.981*** ***(0.222)***	2.627*** *(0.784)*	1.960*** (*0.462*)	2.043*** *(0.580)*	1.695*** *(0.506)*
β^{D}	2.011*** *(0.441)*	2.041*** ***(0.186)***	1.627*** ***(0.289)***	2.124*** ***(0.156)***	2.031*** *(0.598)*	0.583 (*0.724*)	1.597*** *(0.569)*	0.703 *(0.891)*
γ^{ND}	0.493 (*0.341*)	1.014*** ***(0.294)***	1.104*** ***(0.308)***	1.104*** ***(0.279)***	1.305** (0.647)	0.897** (*0.352*)	0.925** *(0.428)*	0.894** *(0.402)*
γ^{D}	1.041** (*0.504*)	0.832** ***(0.339)***	0.340 ***(0.584)***	0.392 ***(0.475)***	1.071* (0.637)	0.384 (*1.133*)	0.795 *(0.967)*	1.068 *(1.028)*
Eq. SE	0.431	0.852	0.847	0.852	0.356	0.341	0.357	0.353
$\bar{R}^2$	0.970	0.947	0.948	0.947	0.974	0.976	0.974	0.974
$\boldsymbol{WT^{\pi}}$	0.052	0.219	0.231	0.479	0.140	0.036	0.276	0.172
$\boldsymbol{WT^{\hat{y}}}$	0.359	0.636	0.231	0.125	0.787	0.676	0.901	0.875

Notes: This table reports the OLS estimates of Equation (1.6) over the sample periods 1979:Q4 – 2008:Q4 and 1987:Q4 – 2008Q4 using four financial stress dummy variables (D^X). D^{CFCI} indicates financial stress related to overall financial conditions, D^{CSPR} denotes credit risk-related stress, D^{SP} identifies stock market bear conditions and D^{LQ} denotes stock market liquidity-related stress. The estimates of β^{ND} and β^{D} denote the policy response to expected inflation in normal times and in times of financial market stress, respectively. The estimates of γ^{ND} and γ^{D} denote the policy response to output gap in normal times and in times of financial market stress, respectively. The last two rows (WT^{π} and $WT^{\hat{y}}$) represent the Wald test p-values of the null hypotheses that $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.17 notes.

Table A1.20: Basic and augmented Taylor rules – financial crisis effect – alternative sample periods

	Panel A: Basic Taylor rule	Panel B: Augmented Taylor rules				Panel C: Basic Taylor rule	Panel D: Augmented Taylor rules			
	1979:Q4 – 2008:Q4					1987:Q4 – 2008:Q4				
	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*	D^{07-08}	*CFCI*	*CSPR*	*SP*	*LQ*
α	-0.329 ***(0.638)***	-0.848 *(0.806)*	-0.129 ***(0.844)***	-0.974 ***(0.617)***	-0.507 ***(0.654)***	-0.312 *(0.925)*	-0.363 *(0.992)*	1.371 *(1.293)*	-0.489 *(1.063)*	0.486 *(1.057)*
ρ_1	0.695*** ***(0.066)***	1.250*** *(0.130)*	0.702*** ***(0.071)***	0.693*** ***(0.061)***	0.703*** ***(0.066)***	1.366*** *(0.091)*	1.431*** *(0.087)*	1.401*** *(0.087)*	1.468*** *(0.099)*	1.445*** *(0.102)*
ρ_2	-	-0.418*** *(0.096)*	-	-	-	-0.509*** *(0.082)*	-0.567*** *(0.080)*	-0.540*** *(0.079)*	-0.604*** *(0.088)*	-0.573*** *(0.090)*
β	-	2.059*** *(0.305)*	2.002*** ***(0.284)***	1.925*** ***(0.180)***	1.914*** ***(0.186)***	-	1.876*** *(0.387)*	1.744*** *(0.367)*	1.784*** *(0.405)*	1.582*** *(0.403)*
γ	-	0.839*** *(0.200)*	0.906*** ***(0.273)***	0.851*** ***(0.239)***	0.991*** *(0.243)*	-	1.067*** *(0.265)*	0.987*** *(0.241)*	0.757** *(0.323)*	1.170*** *(0.267)*
μ^{ND}	-	0.593*** *(0.193)*	-0.546 ***(1.150)***	0.042*** ***(0.014)***	-0.040 ***(0.042)***	-	0.522** *(0.219)*	-1.599 *(1.107)*	0.032* *(0.019)*	0.029 *(0.053)*
μ^{D}	-	2.257*** *(0.789)*	-2.648*** *(0.947)*	0.136*** ***(0.028)***	0.359*** ***(0.076)***	-	2.467** *(1.026)*	-3.962*** *(0.935)*	0.210** *(0.080)*	0.569*** *(0.170)*
β^{ND}	1.904*** ***(0.182)***	-	-	-	-	1.625*** *(0.358)*	-	-	-	-
β^{D}	-0.321 ***(0.466)***	-	-	-	-	-1.595 *(1.197)*	-	-	-	-
γ^{ND}	1.032*** ***(0.257)***	-	-	-	-	1.236*** *(0.260)*	-	-	-	-
γ^{D}	1.152*** ***(0.196)***	-	-	-	-	2.316*** *(0.728)*	-	-	-	-
Eq. SE	0.809	0.397	0.817	0.817	0.822	0.313	0.319	0.309	0.331	0.311
$\bar{R}^2$	0.952	0.975	0.951	0.951	0.951	0.980	0.979	0.980	0.977	0.980
$\boldsymbol{WT^X}$	-	0.037	0.007	0.04	0.000	-	0.064	0.067	0.038	0.002
$\boldsymbol{WT^\pi}$	0.000	-	-	-	-	0.004	-	-	-	-
$\boldsymbol{WT^{\hat{y}}}$	0.645	-	-	-	-	0.239	-	-	-	-

Notes: This table reports the OLS estimates of Equation (1.6) in Panel A and C and Equation (1.5) in Panel B and D over the sample periods 1979:Q1 – 2008:Q4 and 1987:Q4 – 2008:Q4 using financial crisis dummy variable ($D^{07\text{-}08}$). $D^{07\text{-}08}$ takes value one during 2007:Q4 – 2008:Q4 and zero otherwise. Estimates of μ^{ND} and μ^{D} denote policy response to a financial indicator before the crisis and during the crisis, respectively. Estimates of β^{ND} and β^{D} denote policy response to expected inflation before the crisis and during the crisis, respectively. Estimates of $\gamma^{ND}$and γ^{D} denote policy response to output gap before the crisis and during the crisis, respectively. The last three rows (WT^X , WT^π and $WT^{\hat{y}}$) represent Wald test p-values of the null hypotheses that $\mu^{ND} = \mu^{D}$, $\beta^{ND} = \beta^{D}$ and $\gamma^{ND} = \gamma^{D}$, respectively. See also Table A1.17 notes.

Figure A1.1. Real-time and *ex post* measures of output gap

Notes: The figure plots output gap measures based on real-time real GDP data (solid line) and *ex post* real GDP data (dashed line) over the sample period 1985:Q1 – 2008:Q4. The output gap is constructed using Hodrick-Prescott filter.

Chapter 1 – Appendix B

This Appendix presents the technical details and basic theoretical framework of the GMM estimation based on Hayashi (2000).

Firstly, a linear regression model can be written as follows:

$$y_t = x_t'\beta + u_t \qquad t = 1,2,3...T \tag{B1.1}$$

where y_t is a dependent variable, x_t is a (p x 1) vector of regressors, β is a (p x 1) vector of parameters and u_t is an unobserved error term. In order to estimate the vector of parameters, it is assumed that there is a set of moment conditions that β should satisfy. Define z_t as a (q x 1) vector of q instruments. The estimation of β is based on the population moment conditions that all instruments are orthogonal to u_t, i.e. the orthogonality condition holds:

$$E\left[z_t\left(y_t - x_t'\beta\right)\right] = E\left[z_t u_t\left(\beta\right)\right] = 0 \tag{B1.2}$$

In addition, the instruments also must be correlated with the regressors. The true value of β is the solution to the system of equations in (B1.2). The order condition of identification states that there should be at least as many instruments as parameters to estimate ($q \geq p$). The parameters are said to be *just-identified* by the population moment conditions if $q = p$ and *over-identified* by the population moment conditions if $q > p$.

The basic idea of the method of moments is to find an estimate for β such that the corresponding sample moment conditions are also equal to zero and solves the q-equation system below:

$$g_T\left(\hat{\beta}\right) = \frac{1}{T}\sum_t z_t\left(y_t - x'\hat{\beta}_t\right) = \frac{1}{T}\sum_t z_t u_t\left(\hat{\beta}\right) = \frac{1}{T}Z'u\left(\hat{\beta}\right) = 0 \tag{B1.3}$$

For $q > p$, the above system of equations may not have an exact solution, i.e. there may not be possible to find a $\hat{\beta}$ that sets all sample moment conditions to zero. Therefore, $\hat{\beta}$ will be chosen so that $g_T\left(\hat{\beta}\right)$ is as close as possible to zero. Then, the GMM objective function is:

$$J\left(\hat{\beta},W_T\right)=T\ g_T\left(\hat{\beta}\right)'W_T\ g_T\left(\hat{\beta}\right) \tag{B1.4}$$

where W_T is a (q x q) weighting matrix used to construct a quadratic form of the moment conditions. The GMM estimator of β is the $\hat{\beta}$ that minimises Equation (B1.4):

$$\hat{\beta}_{GMM}=\arg\min_{\hat{\beta}} J\left(\hat{\beta},W_T\right) \tag{B1.5}$$

Following Clarida, Gali and Gertler (1998), an estimable basic Taylor rule specification is written as follows:

$$i_t=(1-\rho)\alpha+\beta E\left[\pi_{t+k}\,|\,\Omega_t\right]+\gamma E\left[\hat{y}_t\,|\,\Omega_t\right]+\rho i_{t-1}+\upsilon_t \tag{B1.6}$$

where i_t represents the policy interest rate that partially adjusts to the target with the degree of smoothing ρ, π_{t+k} is the rate of inflation between t and t + k, the contemporaneous output gap is denoted by $\hat{y}_t$, E is the expectations operator, Ω_t represents the information available to policymakers at the time of decision making, and υ_t is an exogenous random shock to the interest rate.

Using the GMM framework, the unobserved forecast values in (B1.6) can be replaced with their actual (realised) values assuming rational expectations:

$$i_t=(1-\rho)\alpha+(1-\rho)\left(\beta\pi_{t+k}+\gamma\hat{y}_t\right)+\rho i_{t-1}+\varepsilon_t \tag{B1.7}$$

where $\varepsilon_t\equiv-(1-\rho)\left\{\beta\left(\pi_{t+k}-E\left[\pi_{t+k}\,|\,\Omega_t\right]\right)+\gamma\left(\hat{y}_t-E\left[\hat{y}_t\,|\,\Omega_t\right]\right)\right\}+\upsilon_t$, i.e. a linear combination of the forecast errors and the exogenous shock to the interest rate. Let z_t ($z_t\in\Omega_t$) be a vector of instrumental variables that are orthogonal to the error term ε_t so that $E\left[\varepsilon_t\,|\,z_t\right]=0$. Thus, the following set of orthogonality conditions must be satisfied:

$$E\left[i_t-(1-\rho)\alpha-(1-\rho)\left(\beta\pi_{t+k}+\gamma\hat{y}_t\right)-\rho i_{t-1}\,|\,z_t\right]=0 \tag{B1.8}$$

Chapter 2: Variance decomposition of US government bond market and the impact of monetary policy

2.1 Introduction

Low and stable inflation along with sustained economic growth define the period of the so-called Great Moderation that started in the mid-1980s. Macroeconomic stability was accompanied by - some argue delivered by - relatively stable, simple and predictable monetary policy conduct. Regrettably, this era of tranquillity was brought to an abrupt end with the global financial crisis in 2007 - 2009. As the zero lower bound (ZLB) on policy rates constrained policymakers around the globe, conventional monetary policy was proved to be powerless to boost aggregate demand. Consequently, the Federal Reserve (Fed) turned to unconventional policies, such as liquidity facilities to improve financial market conditions and quantitative easing (QE), i.e. outright purchases of Treasury bonds and other similar assets from the private sector, to reduce longer-term interest rates and to increase aggregate demand. Within six years, the balance sheet of the Fed underwent an unprecedented expansion until the end of QE was announced in October 2014. The first increase in the federal funds rate (FFR) target in nearly a decade followed in December 2015, indicating the beginning of going back to normal. Understanding the wider asset market impact of this monetary strategy is crucial for policymakers and particularly important is the relationship between monetary policy uncertainty and the bond market developments.

This chapter investigates the sources of variation in US government bond returns and the role of monetary policy over the last three decades. Two strands of the bond market literature are relevant for the empirical analysis here. The first strand includes studies that assess the role of macroeconomic forces, most importantly inflation, in determining developments in the term structure of interest rates. As Duffee (2015) notes, the significance of a model-implied inflation risk for nominal bonds within term structure models varies considerably from very high (Piazzesi and Schneider; 2007; Bansal and Shaliastovich, 2013) to relatively low (Wachter, 2006). As an alternative to theoretical models, another common approach to determine asset returns in terms of macroeconomic forces is the log-linear approximation to the standard present value framework used in a combination with a reduced-form vector autoregressive model (VAR) in the spirit of Campbell and Shiller (1988). The decomposition of bond returns to the revisions in expectations ("news") about future excess returns, inflation and real interest rates was pioneered by Campbell and Ammer (1993). Using this approach, it is commonly found that inflation news explains most of the variance in long-term government bond returns in the US (Campbell and Ammer, 1993; Engsted and Tanggaard, 2007) as well as in other countries (Barr and Pesaran, 1997; Cenedese and Malluci, 2016).

The second strand of the literature examines the impact of monetary policy on the term structure of interest rates. With respect to conventional monetary policy, the evidence shows that Treasury yields across maturities respond positively and significantly to an exogenous increase in the FFR, nevertheless, the magnitude of a response tends to diminish at longer maturities (Kuttner, 2001; Cochrane and Piazzesi, 2002; Gurkaynak, Sack, and Swanson, 2005). Following the implementation of QE, there has been a surge of studies that examine its impact on bond yields. Using various econometric techniques, it is generally found that QE was effective in reducing long-term Treasury bond yields. The existing literature identifies two key channels of transmission that explain the decline in long-term yields: the signalling and portfolio balance channels. According to the signalling channel, QE leads to lower expectations about future short-term interest rates that through the expectations theory of the term structure result in lower long-term rates (Christensen and Rudebusch, 2012; Bauer and Rudebusch, 2014). On the other hand, the portfolio balance channel implies that the decline in the supply of long-term bonds in the market compresses the term (risk) premium and, thus, reduces their yields (Gagnon et al., 2011; D'Amico et al., 2012). Nevertheless, the empirical evidence as to which channel is more important is rather mixed, indicating that the understanding of how QE led to lower bond yields is still incomplete.

This chapter takes an alternative approach to estimate the bond market response to monetary policy and to identify the sources of this response. Specifically, the extension of Campbell and Ammer's (1993) framework suggested by Bernanke and Kuttner (2005) is applied to bond market returns with respect to both conventional and unconventional monetary policy.[57] Firstly, unexpected current period excess returns on the 2-, 5- and 10-year Treasury bonds are decomposed to news about future excess returns, inflation and real interest rates. Secondly, the impact of conventional and unconventional monetary policy on

[57] The study that is close to the analysis in this chapter is that by Bredin, Hyde and O'Reilly (2010). They employ a similar approach to examine the pre-crisis (1994-2004) conventional monetary policy impact on domestic and international bond markets in the US, UK and Germany. In the case of the US, they do not find significant effects.

Treasury bond returns and their components is examined. The sample period commences during the Great Moderation and ends in the aftermath of the recent financial crisis (1985:1 – 2014:2). In order to capture conventional policy shifts, the federal funds rate-based measures are used, while unconventional policies are captured using changes in the monetary base. The use of quantity-based indicators is motivated by a number of recent studies that evaluate the role of the monetary base, or the supply of reserves, as an alternative operating target for monetary policy (Curdia and Woodford, 2011; Gertler and Karadi, 2013). Thus, the contribution of the analysis in this chapter is three-fold. Firstly, the empirical approach allows explaining the bond market reaction to monetary policy changes in terms of revisions in expectations about macro-fundamentals, such as real interest rates and inflation, and future expected excess bond returns, i.e. the risk (term) premium. This set-up has not been widely applied for the bond market. The sample period covers the last three decades including the global financial crisis. Secondly, special attention is paid to the role of the financial crisis and unconventional policies subsequently adopted by the Fed. This is the first attempt to analyse the effects of quantitative easing within the VAR-based returns variance decomposition framework. Finally, shorter maturities are also considered in addition to the commonly analysed 10-year Treasury bonds. Thus, it is possible to compare the effects across the yield curve.

The main results can be summarised as follows. Across the maturities, news about future inflation is the key factor in explaining the variance in unexpected excess Treasury bond returns during the sample period. Meanwhile, the role of news about expected excess bond returns and real interest rate news is typically much less relevant. Regarding the effect of conventional and unconventional monetary policy actions, monetary easing typically leads to higher unexpected excess bond returns**.** Nevertheless, the bond market response to conventional policy has grown somewhat weaker since the early 1990s. This may reflect changes in the way that the Fed implements and communicates its monetary policy decisions. With respect to quantity-based monetary policy indicators, the results are largely driven by the peak of the financial crisis in autumn 2008 when unprecedented expansion in the Fed's balance sheet was accompanied by a stronger bond market response to money growth. Furthermore, the findings highlight the importance of inflation news in explaining the bond market reaction to monetary policy. The positive effect of monetary easing on unexpected excess returns mainly comes from a corresponding negative effect on inflation expectations. Thus, the evidence is overall not supportive for the portfolio balance mechanism's prediction of a strong role for the risk (term) premium to explain bond market reaction to QE policy. The results are reasonably robust to various sensitivity checks, related to the specification of the underlying VARs and monetary policy proxies.

The chapter has the following structure. The review of relevant literature is presented in Sections 2.2, 2.3. and 2.4. Section 2.5 explains the methodology. Section 2.6 describes the dataset and explains the proxies used to identify monetary policy changes. Section 2.7 contains the empirical results from the main analysis, while Section 2.8 discusses the robustness analysis. Section 2.9 concludes.

2.2 Bond market determinants

This and the next sections provide the review of related studies from the two strands of the literature that are relevant for this chapter. The survey starts with the overview of studies that assess the role of macroeconomic forces, most importantly inflation, in determining the term structure of interest rates. The review in the following sections is focused on empirical studies investigating the impact of conventional and unconventional monetary policy by the Federal Reserve on market interest rates.

In general, yield curve dynamics are influenced by expectations about short-term nominal interest rates, i.e. the sum of short-term real rates and inflation, and the term (risk) premium (Campbell, Pflueger and Viceira, 2015). A vast amount of literature is focused on the models of the term structure of interest rates to capture dynamics in the yield curve. Several macroeconomic factors are distinguished as the potential determinants of bond yields. These factors range from inflation, real activity, and consumption (Ang and Piazzessi, 2003; Piazzesi and Schneider; 2007; Chernov and Mueller; 2012; Joslin, Priebsch and Singleton, 2014) to monetary policy (Ang et al., 2011). Campbell, Pflueger and Viceira (2015) build a general equilibrium model for asset pricing and investigate how macroeconomic shocks and changes in monetary policy affect the bond risk premium. Inflation factor has received a lot of attention in the literature. However, Duffee (2015) notes that the model-implied relative importance of inflation risk in explaining yield curve volatility varies substantially across these models. With respect to standard dynamic term structure models, inflation risk accounts for almost entire variation in nominal yields (Piazzesi and Schneider, 2007; Bansal and Shaliastovich, 2013). On the other hand, models with habit formation preferences imply a much smaller role of inflation expectations for the variance of bond yields (Wachter, 2006).

Alternatively, a common approach to determine asset returns in the empirical finance literature is to use the combination of the log-linear approximation to the standard present value framework and a reduced-form VAR model in the spirit of Campbell and Shiller (1988). This method does not rely upon strong theoretical assumptions and uses accounting identities to link unexpected excess returns on assets to revisions in rational expectations ("news") about the components of these returns. Since its origination, this methodology has been widely applied to stock market returns (Campbell, 1991; Campbell and Ammer, 1993; Ammer and Mei, 1996; Vuolteenaho, 2002; Engsted and Tanggaard, 2004; Bernanke and Kuttner, 2005; Bredin et al., 2007; Botshekan, Kraeussl, and Lucas, 2012; Garrett and Priestly, 2012; Maio, 2014).

In one of the initial studies, Campbell (1991) decomposes unexpected real returns on the US stock market into two components: the discounted sum of revisions in rational expectations regarding future dividend flows (*"cash flow news"*) and the discounted sum of revisions in expectations regarding future stock returns (*"discount rate news"*). Alternatively, he also proposes a three-way decomposition for stock returns in excess of a risk-free short-term interest rate. In this case, unexpected excess stock returns are explained in terms of cash flow news, discount rate news and the discounted sum of revisions in expectations about future real short-term interest rate (*"real interest rate news"*). Consequently, the variance of unexpected returns can be written in terms of the variance of the components and the covariance terms between them. A vector autoregression model can then be used to obtain empirical proxies for the news components from reduced-form residuals. Campbell (1991) finds that the variance of the cash flow news component explains approximately a third of the total variance of stock returns, while the variance of discount rate news typically accounts for the major part of it. With respect to the US, the dominance of discount rate news in determining stock returns is also highlighted in the subsequent literature (Campbell and Ammer, 1993; Engsted and Tanggaard, 2004; Bernanke and Kuttner, 2005). On the other hand, several recent studies point out that the role of cash flow news may be understated (Garrett and Priestly, 2012; Maio, 2014).

The VAR-based returns variance decomposition is extended to the bond market by Campbell and Ammer (1993).[58] They analyse returns on the US stock market and nominal zero-coupon US government bonds with 10-year maturity over the period 1952:1 – 1987:2. Unexpected excess bond returns are decomposed into the *"inflation news"* component, i.e. the sum of revisions in expectations about future inflation, the *"real interest rate news"*

[58] This methodology has also been applied to assets other than stocks and government bonds. For instance, Nozawa (2014) decomposes corporate credit spreads into the expected credit losses innovations and expected returns news components. Several studies employ the variance decomposition of asset returns for the analysis of housing or real estate markets (Bredin, O'Reilly, and Stevenson, 2011; Engsted and Pedersen, 2014).

and *"risk premium news"* components, i.e. the sums of revisions in expectations about future real rate and future excess bond returns, respectively.[59] With respect to the bond market, the results indicate that inflation news is the key driving force of the overall variance of unexpected excess returns or at least as important determinant as risk premium news. On the other hand, real interest rates do not play any role in explaining excess bond returns (Campbell and Ammer, 1993).

Subsequently, a large number of studies used the approach by Campbell and Ammer (1993) to examine bond market developments in the US and other countries (Barr and Pesaran, 1997; Engsted and Tanggaard, 2001; Valckx, 2004; Engsted and Tanggaard, 2007; Bredin, Hyde and O'Reilly, 2010; Breedon, 2012; Nozawa, 2014; Cenedese and Mallucci, 2016). Typically, inflation news is found to be the dominant component of excess bond returns, whilst real interest rate news is generally irrelevant. For instance, Engsted and Tanggaard (2007) estimate a multi-country VAR model to analyse the co-movement of US and Germany government bond markets for the period 1975:7 – 2003:2 based on the cross-correlation between the components of returns across countries. The positive and strong correlation between bond returns is explained by the fact that inflation news component is the key driver of unexpected returns in both countries (Engsted and Tanggaard, 2007).

Similarly, Engsted and Tanggaard (2001) analyse the co-movement of the Danish stock and government bond markets in 1922 – 1996. The results of variance decomposition confirm that the dominant component in explaining variability in the government bond market is inflation news, while stock returns are largely determined by news about future cash flows. Despite the positive relationship between actual returns on stocks and bonds, the findings also indicate that news about future excess returns in the two markets are negatively correlated. Possibly, stock and bond returns respond to different information or to the same information but in a different manner (Engsted and Tanggaard, 2001).

Bredin, Hyde and O'Reilly (2010) analyse bond market movements in the US, UK and Germany. They estimate a multi-country VAR model and conduct the variance decomposition of unexpected excess bond returns for each country in the period 1975:2 – 2004:12. The results indicate that inflation news variance accounts for the major part of the total variability in excess bond returns for all three countries. In line with other studies, they do not find a strong role for expected excess bond returns in determining bond market developments and the share of returns volatility explained by real rate news is negligible.

[59] The risk premium associated with holding bonds is the term premium in the case of government bonds (Campbell and Ammer, 1993; Engsted and Tanggaard, 2007).

Barr and Pesaran (1997) examine the variability in excess returns on nominal and index-linked (real) bonds in the UK over the period 1983:4 – 1993:10. They find that revisions in expected inflation are clearly the most important component of unexpected returns when nominal bonds are considered. Nevertheless, the risk premium also accounts for a considerable share of the total variance of returns. In the case of index-linked bonds, risk premium news becomes much more relevant and inflation news only accounts for a small share of returns volatility. Interestingly, the real interest rate news component is not important for either type of bonds. With respect to relative returns, over 90% of variance in relative returns is explained through revisions in expectations about future inflation, while news about future expected excess returns only plays a small part (Barr and Pesaran, 1997).

The recent study by Cenedese and Mallucci (2016) provides some insight into the variance of bond market returns for thirty one countries, including advanced and emerging market economies. The results for the sample period 2004:1 – 2013:12 indicate that revisions in future expected inflation are the main source of variation in bond market returns for both groups of countries and for all countries taken together. On the other hand, other two news components of unexpected bond returns are not found to be important.

The vast majority of empirical studies report that revisions in inflation expectations determine bond returns to a great extent. Nevertheless, several studies find the opposite result. For the period 1975:1 – 2013:7, Nitschka (2014) provides the evidence that the most relevant determinant of the variability in US bond returns is news about future excess returns. Similarly, Valckx (2004) demonstrates that inflation news may not be the key factor driving the volatility of bond returns in the US during 1954:6 – 2000:12. At least over the long horizon, news about expected future returns appears to be a more important factor.

2.3 Conventional monetary policy effects on market interest rates

Generally, monetary policy tightening increases interest rates at the short-end of the yield curve as the supply of credit tightens. At the long-end, the overall effect is much less clear. It depends on expectations about future short-term interest rates, according to the expectations theory of the term structure, and on changes in expected inflation and real long-term rates (Rolley and Sellon, 1995; Estrella and Mishkin, 1997). In response to a monetary policy shock, long-term interest rates typically change to the same direction as short-term rates, but to a smaller degree. Nevertheless, the direction of a change and the degree of responsiveness of long-term rates may differ depending on market expectations about future monetary policy and the perceived persistence of policy actions (Rolley and Sellon, 1995). For instance, if policy actions are seen as relatively permanent or as the first in a series of future actions, the response of long-term rates may even be greater than the response of short-term rates. Equivalently, if policy actions are expected to be fully offset and reversed in the future, longer-term rates may fall in response to monetary policy tightening.

2.3.1 Early evidence: money and market interest rates

Initial studies investigating the impact of US monetary policy on market interest rates typically use money supply growth, actual and unexpected, as a measure of policy instrument. According to Gibson (1970), the response of interest rates to changes in money supply can be explained through the effects of liquidity, income, and inflationary (price) expectations. The liquidity-preference relationship, i.e. the negative relationship between the quantity of money demanded and interest rates, forms the basis for the liquidity effect. An increase in money stock, which must be held by someone, induces a shift towards other assets and to a fall in yields on these assets as their prices increase.[60] Thus, money growth should lower interest rates. Alternatively, the inflationary expectations effect implies higher interest rates due to higher expected price level following expansion in money supply. Finally, if nominal income increases at the same time as money stock growth accelerates, this leads to higher demand for money. Consequently, it puts upward pressure on interest rates counteracting the negative liquidity effect. According to the income effect, the net impact of higher money supply on interest rates may then be very small or even zero if the liquidity effect is fully offset.

In order to test for these effects, Gibson (1970) models short- and long-term interest rates as the functions of current and lagged values of monetary aggregates. The findings indicate a significant and negative initial effect of an increase in money stock on a short-term interest rate. Nevertheless, it is offset by the positive income effect within several months. Also, the initial liquidity effect is found to be insignificant in the case of long-term interest rates, i.e. yields on government and corporate bonds.

The early findings of the liquidity effect are challenged by subsequent studies. Reichenstein (1987) provides a brief survey of the empirical evidence with respect to

[60] Money demand is assumed to be constant. The yields decline until the new equilibrium is reached where the new money supply is equal to money demand.

money supply effects on short-term interest rates. He argues that, at least since 1975, the evidence is generally not supportive of the liquidity effect. On the one hand, this could simply indicate that the inflationary expectations effect is stronger than the liquidity effect. In addition, Reichenstein (1987) emphasises the policy anticipation effect that implies the tendency of market participants to expect corrective actions by the Fed in the near future in response to higher than expected money stock growth. Since both effects tend to push up interest rates, the liquidity effect, even if present in the short-term, is likely to be fully offset relatively quickly. Given rather mixed evidence of the impact of money supply changes on interest rates, Reichenstein (1987) concludes that the Fed has little influence over short-term interest rates.[61]

With respect to long-term interest rates, early empirical evidence using quantity-based monetary policy measures is also rather mixed (Akhtar, 1995). For instance, Feldstein and Chamberlain (1973) find a significant and negative effect of monetary base growth on a long-term yield. Also, Cochrane (1989) argues in favour of the liquidity effect and demonstrates that there is a significant and negative correlation between money supply growth and both the 3-month Treasury bill rate and 20-year bond yield. On the contrary, Mishkin (1981) shows that unexpected changes in money supply generally have no significant impact on bond returns.

2.3.2 Federal funds rate target and market interest rates

The empirical evidence is more consistent across studies if the federal funds rate target is used as a monetary policy instrument. The FFR was set as the monetary policy target for the first time from 1972 through 1979. Thereafter, the Fed implemented the non-borrowed and borrowed reserves operating procedures over the periods 1979 – 1982 and 1982 – 1988, respectively. Since the late 1980s, the federal funds rate targeting has been in operation and this procedure continues to be implemented today (Strongin, 1995; Walsh, 2003). Accordingly, it is a standard practice in the literature to use the funds rate as a proxy for monetary policy actions over this period. Two empirical approaches are predominantly applied to estimate monetary policy effects on the term structure of interest rates, which also account for the endogeneity problem: event studies based on higher-frequency data

[61] Akhtar (1995) notes that when using narrow quantity-based variables to measure monetary policy shocks, it is typically found that expansionary policy leads to a decline in short-term interest rates. On the other hand, broader monetary aggregates, such as M1 or M2, tend to show that monetary expansion does not result in lower short-term interest rates.

and vector autoregression models typically using lower-frequency data.[62] The evidence with respect to both approaches is discussed in the subsequent paragraphs.

The prominent study of Cook and Hahn (1989) is the first one to examine the response of market interest rates to changes in the federal funds rate target. For the period 1974:9 – 1979:9, the regression analysis over the days of changes in the target rate reveals a positive and significant impact of funds rate target changes on the term structure of interest rates. In response to a 1-percentage-point increase in the target rate, the 3-month and 1-year Treasury bill rates rise by 55 and 50 basis points, respectively. Although the effect becomes smaller towards the long-end of yield curve, the effect remains statistically significant and the 5-, 10- and 20-year Treasury yields increase by 21, 13 and 10 basis points, respectively.

Nevertheless, subsequent event studies examining later sample periods fail to find strong evidence of a significant monetary policy impact on long-term rates. For the period 1987:10 – 1995:7, Rolley and Sellon (1995) show that following a 1-percentage-point rise in the FFR target the 1-year and 30-year Treasury yields increase by 22 and 4 basis points, respectively. Furthermore, the response of the long-term rate is statistically insignificant. Kuttner (2001) revisits the findings by Cook and Hahn (1989) for the later sample period spanning 1989:6 – 2000:2. The magnitude of interest rate responses is smaller across maturities than in the earlier period as in the original study. Furthermore, there is no significant monetary policy effect on market interest rates with maturities beyond 5 years.

Rolley and Sellon (1995) note that markets tend to anticipate policy actions in advance and market rates move prior to actual changes in the funds rate target. As Kuttner (2001) explains, monetary policy actions have become much less of a surprise to market participants in more recent period. To address this anticipation effect, he constructs a measure of the unexpected component of a change in the funds rate target using data for the federal funds rate futures. Subsequently, Kuttner (2001) investigates the effects of both expected and unexpected changes in the target rate on market interest rates for the period 1989:6 – 2000:2. First of all, an expected rate change is found to have no significant impact. In contrast, the unexpected component has a positive and significant effect on all interest rates under consideration. Secondly, the response tends to gradually decline with the maturity. For instance, a 1-percentage-point positive shock to the funds rate target is

[62] The endogeneity issue may arise due to a monetary policy response to developments in financial markets or a simultaneous reaction of both to other news (Rigobon and Sack, 2004). Alternatively, it may also arise if policy actions reveal some private information that the Fed possesses about future economic developments (Romer and Romer, 2000).

associated with an increase in the 3-month Treasury bill rate and 30-year bond yield by 79 and 19 basis points, respectively.

Ever since Kuttner (2001), the focus in the literature has shifted towards the effects of unexpected monetary policy actions. Poole, Rasche and Thornton (2002) argue that monetary policy surprises as in Kuttner (2001) contain a measurement error due to other news that potentially cause movements in the federal funds futures market even outside monetary policy events or headline news days. They account for this potential bias in the estimated interest rate response using the errors-in-variables estimator. Nevertheless, the results remain largely in line with the previous evidence. Overall, an unexpected change in the FFR target has a significant and positive impact on interest rates across the maturities of three months to thirty years. Also, the magnitude of the effect declines as maturity lengthens. The sub-sample analysis shows that after February 1994, when changes in the Fed's communication practice took place, longer-term yields become somewhat less responsive to policy surprises. Possibly, this may be the consequence of greater transparency and better predictability of monetary policy conduct (Poole, Rasche and Thornton, 2002).

As Gurkaynak, Sack and Swanson (2005) demonstrate, a single factor, i.e. changes in the current target rate, may not be enough to adequately capture monetary policy surprises. They identify two factors, target and path, that determine the effects of monetary policy on the 2-, 5- and 10-year Treasury yields in 1990:1 – 2004:12. The target factor represents unexpected changes in the FFR target and is measured using tick-by-tick data on the current-month federal funds futures contract rate. The path factor is associated with the expectations about future monetary policy over the next one year and is estimated using Eurodollar futures rates. The results show that the target factor has a positive and statistically significant effect on Treasury yields that declines in magnitude towards the long-end of the term structure. Furthermore, yields also respond positively and significantly to the path factor that appears to be even more important than the target factor in the case of the 5- and 10-year rates.

With respect to the federal funds futures data and the identification of monetary policy shocks, Hamilton (2008) proposes the generalisation of formulas suggested in the literature. The approach takes into account the deviations of the effective federal funds rate from its target and does not condition on the days of target changes. Despite this new methodology, the results, obtained for the period 1988:10 – 2006:12, are consistent with Kuttner (2001) and Pool, Rasche and Thornton (2002). The unexpected federal funds rate target change has a positive and significant effect on both short- and longer-term interest rates, with long-term yields being less responsive.

On the other hand, Thornton (2014) argues that empirical studies using market-based measures as proxies for monetary policy shocks do not account for the *"joint-response bias"*. In addition to monetary policy news, both market-based measures of monetary policy surprises and interest rates are influenced by other news. Thus, Thornton (2014) develops the methodology that corrects for the bias in the estimated interest rate responses. The impact of the federal funds target rate on market interest rates is much smaller and less significant than reported in earlier studies after the joint-response is accounted for.[63]

In addition to event studies, several other methods have been applied in empirical studies to estimate the monetary policy impact on market interest rates. For instance, Rigobon and Sack (2004) achieve the identification of shocks through the heteroscedasticity that is present in daily data. This methodology mainly requires the assumption that the variance of monetary policy shocks is higher on the days of the FOMC meetings and/or other relevant policy events. Hence, the shift in variance is enough to trace down asset price response to policy shocks. The results are largely consistent with the evidence from event studies. For the sample period 1994:1 – 2001:11, the yields on Treasuries with 6-month, 1-, 2-, 5-, 10-, and 30-year maturities increase significantly in response to a contractionary policy shock. Also, the response is found to be declining in magnitude with the maturity.

Another strand of the literature utilises a structural VAR framework. For instance, Edelberg and Marshall (1996) investigate the effects of monetary policy on government bond yields with 1-month to 15-year maturity in the period 1947 - 1995. The estimated impulse response functions show a significant and positive effect of an exogenous increase in the federal funds rate on short-term interest rates. In line with the event-study evidence, the response of a yield declines considerably and becomes much less significant as maturity increases above one year. Similarly, Evans and Marshall (1998) estimate three VAR models using alternative identification schemes in order to examine the effects of exogenous monetary policy shocks on nominal interest rates in 1965:1 – 1995:12. The findings from all three models provide support for the liquidity effect. A contractionary monetary policy shock induces a significant, although temporary, increase across the short

[63] Furthermore, the estimated response of Treasury rates to monetary policy surprises is negative in the period since early 2000s and, for longer maturities, it is also statistically significant. One possible explanation for this finding is that monetary policy tightening tends to raise real interest rates due to lower inflation expectations (Thornton, 2014).

end of the term structure. On the other hand, longer-term rates, such as the yields on 3- and 10-year bonds, are almost unaffected by monetary policy tightening.

As noted by Berument and Froyen (2009), strong monetary policy effects with respect to long-term market interest rates are typically found using a single-equation approach, such as an event study. They provide the comparative study of two methodologies to estimate monetary policy effects on longer-term interest rates. The VAR-based results indicate a rather mild positive reaction of the 1-year interest rate to an increase in the federal funds rate. The corresponding estimate for the 10-year rate is much smaller. Also, the effects of monetary policy on both interest rates become even smaller and insignificant in the period since the mid-1980s. On the contrary, the results from single-equation estimations show a positive and significant response of both interest rates to an unexpected change in the federal funds target rate.

On the other hand, several recent VAR-based studies confirm the strong effect of monetary policy shocks on longer-term yields as they take alternative identification approaches. For instance, Beckworth, Moon and Toles (2012) estimate the structural VAR model that is identified long-run monetary neutrality restrictions. For the period 1979:10 - 2007:12, the results imply that an expansionary monetary policy shock, identified through innovations in the monetary base, leads to significantly lower interest rates across all maturities considered. Moreover, the magnitude of the response does not appear to decline sharply with the maturity. In addition, Gertler and Karadi (2015) employ high-frequency measures of monetary policy surprises in a low-frequency VAR model. Their identification method also takes into account the effects of the forward guidance of monetary policy. For the period 1979:7 – 2012:6, the impulse response analysis shows that a contractionary monetary policy shock significantly increases nominal rates on government bonds with 1-, 2-, 5-, and 10-year maturities, albeit longer-term rates are affected to a smaller degree.

2.3.3 Monetary policy and bond returns

Alternatively, one could analyse the effects of monetary policy on bond returns instead of bond yields. The price of a bond is inversely related to its yield, thus, tighter monetary policy stance is expected to decrease bond prices and, in turn, bond returns. For instance, Johnson et al. (2003) examine returns on government and corporate bond indices for the period 1973:1 – 1999:6. They distinguish between the restrictive and expansionary monetary policy regimes and find that an average return on an index is greater during expansionary monetary policy periods as compared to an average return in restrictive monetary conditions.

An alternative and attractive approach to analyse bond returns reaction to monetary policy is based on the returns variance decomposition developed by Campbell and Ammer (1993). This framework is taken one step further by Bredin, Hyde and O'Reilly (2010). In the spirit of Bernanke and Kuttner (2005), they examine the response of unexpected government bond returns and three components of returns to monetary policy surprises in the US, UK and Germany.[64] For the period 1994:2 – 2004:12, they find no significant impact of US monetary policy on the 10-year Treasury bond returns. Similarly, monetary policy shock has no significant effect on news about inflation, real interest rate or future excess bond returns. Also, there is no evidence of the spillover effect since foreign bond returns do not appear to respond to the Fed's actions. With respect to Germany, a positive effect on current excess bond returns of surprise domestic monetary policy tightening is explained by significant downward revisions in expectations regarding future inflation (Bredin, Hyde, and Reilly, 2010). For the UK, they find a significant and negative impact of a contractionary domestic monetary policy shock on domestic bond returns that is mainly due to upward revisions in inflation expectations. This contrasting effect of monetary policy tightening on inflation expectations in the UK and Germany may be explained by the degree of credibility of the central bank in each country. With respect to international spillovers, the evidence also implies that monetary policy in the UK has some significant impact on the German bond market.[65]

2.4 Unconventional monetary policy effects on market interest rates

2.4.1 The Fed's response to the crisis: a short overview

Several major events caused havoc in global financial markets in September 2008. This had severe consequences on the functioning of financial markets and liquidity conditions, and triggered a prompt response by policymakers at the Fed and other central

[64] Bernanke and Kuttner (2005) extend the VAR-based methodology of returns variance decomposition by adding a market-based measure of monetary policy shock as an exogenous variable in the estimated VAR model. This allows estimating the impact of monetary policy shocks on unexpected excess stock returns and three news components of these returns. However, they do not consider the bond market.

[65] Another related study of US monetary policy effects on bond returns and their respective components is provided by Valckx (2004). Nevertheless, monetary policy is not the main focus of the study; therefore, the discussion is not well-developed. For the period 1954:6 – 2000:12, an unexpected increase in the discount rate significantly lowers current unexpected excess bond returns. This effect is mainly accounted for by a positive and significant impact on expected future inflation and future excess returns. On the other hand, an increase in real money supply appears to lower current excess returns.

banks. Between September and November, the Fed announced various changes to existing liquidity facilities and introduced new facilities to alleviate liquidity strains.[66] In addition, central banks in the US, Canada, the UK, Switzerland, Sweden and the euro area simultaneously cut their policy rates by 50 basis points on 8 October.

Nevertheless, the enhanced liquidity provision and complementary monetary policy actions were not sufficient given weak economic outlook at that time. Consequently, the first round of quantitative easing (QE1) was announced on 25 November 2008. In order to revive the housing market, the Fed committed to purchase up to $100 billion worth of direct obligations of the housing-related government-sponsored enterprises (GSEs) and up to $500 billion of mortgage-backed securities backed by Fannie Mae, Freddie Mac and Ginnie Mae. This is also known as the first round of large-scale asset purchases (LSAPs).[67] In addition to setting the federal funds target rate within the range of 0 - 0.25% on 16 December, the Fed also introduced the explicit forward guidance about a future path of its policy rate. It was announced that the FOMC anticipated “exceptionally low levels of the federal funds rate for some time” due to weak economic conditions. The language was later updated on 18 March 2009 to indicate that exceptionally low funds rate target is to prevail for “an extended period”. At the same time, QE1 was extended by announcing further purchases of the MBS and agency debt worth $750 billion and $100 billion, respectively. In addition, the Fed introduced purchases of $300 billion of longer-term Treasury securities.[68]

By the summer of 2010, the economic recovery faltered again, potentially due to intensifying sovereign debt problems in the euro area. In response to high unemployment and low inflation, the new round of quantitative easing (QE2) was announced on 3 November 2010. It included an additional $600bn purchases of longer-term US Treasuries spread out until the end of the second quarter in 2011. Moreover, the Fed’s forward guidance about the policy rate path was updated again in the FOMC statement on 9 August 2011: the federal funds rate target was to remain at low levels “at least through mid-2013”. In order to put further downward pressure on longer-term interest rates, the Maturity Extension Program (MEP) was announced on 21 September 2011. It involved swapping $400 billion worth of Treasuries with the remaining maturity of three years or less for the Treasuries with the remaining maturity of six to thirty years.

[66] Federal Reserve Bank of St Louis provides the timeline of the crisis-related events and policy actions: https://www.stlouisfed.org/financial-crisis/full-timeline#2008.
[67] In this chapter terms LSAP and quantitative easing (QE) are used interchangeably.
[68] The first round of LSAPs was completed by the end of March 2010.

Finally, the third round of quantitative easing (QE3) was announced on 13 September 2012 amid increasing concerns about elevated unemployment rate and weak labour market conditions. It was then decided to purchase \$40bn of MBS every month and the language regarding the funds rate target was updated to signal an exceptionally low level of the target rate "at least through mid-2015". On 12 December 2012, QE3 was expanded to add monthly purchases of longer-term US Treasuries worth \$45 billion. In addition, the date-based forward guidance was replaced by the forward guidance subject to economic thresholds. The Fed announced that an exceptionally low interest rate would prevail as long as the unemployment rate remained above 6.5%, inflation projections for between one to two years ahead did not exceed 2.5% and long-run inflation expectations remained well anchored. Overall, these policies were implemented to "maintain downward pressure on longer-term interest rates, support mortgage markets, and help to make broader financial conditions more accommodative".[69]

Eventually, as the economic and financial conditions continuously improved, the monthly purchases were gradually reduced and eventually terminated in October 2014. As the result of asset purchases financed by central bank money, the total assets of the Fed expanded from \$869 billion in August 2007 to nearly \$4.5 trillion. The end of QE3 was followed by the increase in the federal funds target rate in December 2015, the first upward change in almost a decade, indicating the beginning of monetary policy normalisation.

2.4.2 QE transmission channels

The literature distinguishes two main transmission channels through which asset purchases by a central bank can potentially reduce longer-term yields on government bonds: the *portfolio balance* channel and *signalling* channel.[70] In order to get a better understanding how each channel works, it is useful to decompose a long-term nominal government bond yield into the average level of expected future short-term risk-free interest rates over the life of a bond and the term premium component (Neely, 2015).[71] The

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[69] See the FOMC minutes and the statement in December 2012: http://www.federalreserve.gov/monetarypolicy/fomccalendars.htm#11655.

[70] Several other channels are also discussed in the literature. For instance, the liquidity or market functioning channel implies that a central bank is able to enhance liquidity and improve market trading conditions by creating additional demand for long-term assets and reducing the liquidity premium (Gagnon et al., 2011). Krishnamurthy and Vissing-Jorgensen (2011) discuss a variety of other channels, such as the prepayment risk and the default risk channels. Nevertheless, in terms of the QE impact on government bond yields, the most important channels are likely to be the portfolio balance and signalling channels.

[71] Mathematically it can be expressed as follows: $y_{t,t+n} = y^*_{t,t+n} + TP_{t,n}$, where $y_{t,t+n}$ is the yield on an n-year bond at time t, $y^*_{t,t+n}$ is the average expected short-term interest rate over n years and $TP_{t,n}$ is the term premium on an n-year bond at time t.

first component reflects expected returns earned when rolling over short-term risk-free investments, whilst the second component represents a required additional expected return for holding the risks related to long-term assets, i.e. the risk premium.[72]

The portfolio balance channel implies that central bank purchases of longer-term assets compresses excess returns required on these assets and on their close substitutes, i.e. asset purchases reduce the risk (term) premium component (Gagnon et al., 2011). The portfolio-balance effect rests upon the preferred-habitat literature that assumes imperfect substitutability between assets across maturities and asset classes, i.e. the market segmentation assumption (Culbertson, 1957; Modigliani and Sutch, 1966). Under this assumption, a yield on a specific maturity is determined by demand and supply shocks associated with that maturity. More recently, Vayanos and Vila (2009) have developed the preferred-habitat model of the term structure of interest rates that includes risk-averse arbitrageurs. The model is useful to analyse the potential effects of quantitative easing on bond yields.

It follows that, contrary to standard term structure models, prices and yields of long-term assets depend on changes in the supply of these assets that is publicly available. There are two mechanisms relevant to the portfolio balance channel (D'Amico et al., 2012). The first one is the so-called *duration risk* channel. As a result of central bank long-term asset purchases, the overall supply of securities with long duration available to the private market is reduced. Consequently, the average market duration risk, which is associated with future developments in interest rates, falls. This leads to lower risk premiums and, in turn, lower yields across the term structure. The second mechanism works through the *local supply* or *scarcity* channel (D'Amico et al., 2012; D'Amico and King, 2013). Asset purchases by a central bank make specific assets, such as long-term Treasuries, scarcer. Due to the mismatch between the demand for and the new (reduced) supply of these assets, upward pressure on their prices will compress yields through lower risk premiums demanded by investors.

It is important to note that the portfolio-balance channel does not only work with respect to the assets being purchased but with respect to other assets that are close substitutes for the assets bought. For instance, provided that central bank money and other assets are not perfect substitutes, the sellers of longer-term assets may want to invest their increased cash holdings to other, possibly more risky, assets that earn positive yields. This way, asset prices and aggregate wealth would increase (Gagnon et al., 2011).

[72] While the term premium is the largest component of the risk premium on government bonds, the credit and liquidity premiums may be a part of the risk premium on other long-term assets, such as MBS or agency debt.

On the contrary, the signalling channel works through changes in the expected future short-term interest rates induced by monetary policy actions, such as the forward guidance and asset purchases. The mechanism is based upon the standard expectations hypothesis of the term structure of interest rates (D'Amico, et al., 2012; Bauer and Neely, 2014). For instance, QE announcements typically signal more accommodative policy stance than otherwise. Thus, it implies a lower future policy rate, i.e. lower expected short-term rates, perhaps due to weaker economic conditions. Moreover, unconventional policy may also imply that a central bank is willing to temporarily change its reaction function or deviate from a normal policy path in order to keep the policy rate unusually low (Bauer and Neely, 2014). Consequently, lower expected short-term interest rates for a longer period lead to a decline in longer-term yields.

2.4.3 Empirical evidence: QE and Treasury yields

The implementation of quantitative easing around the globe has sparked a boom of empirical studies evaluating short-term and longer-term effects of such unconventional monetary policies. The literature is largely focused on the QE effects with respect to longer-term Treasury bonds and other long-term interest rates (Gagnon et al., 2011; Christensen and Rudebusch, 2012; Hamilton and Wu, 2012; Thornton, 2012; D'Amico and King, 2013). On the other hand, a number of studies also look into the effects on other asset prices, such as stocks and exchange rates (Wright, 2012; Rogers, Scotti and Wright, 2014), and macroeconomic variables (Chung et al., 2012; Baumeister and Benati, 2013; Gambacorta et al., 2014). Others examine the international effects of unconventional monetary policy (Chen et al., 2015; Neely, 2015). Another strand of the literature sheds some light on the forward guidance effects at the ZLB (Campbell et al., 2012; Moessner, 2013; Raskin, 2013). Nevertheless, it is inherently difficult to accurately measure the unexpected component of quantitative easing policy.[73] Thus, the majority of the empirical studies analysing the response of bond yields to asset purchases rely on the event-study approach. As discussed in Bauer and Neely (2014), it is also not an easy task to separate the effects of the forward guidance from the effects of asset purchase announcements as the policy statements tend to contain both types of news.[74] This section provides a brief

[73] Martin and Milas (2012) provides a brief summary of studies evaluating the effects of the quantitative easing programmes in the US and UK. They also discuss econometric methods typically used and outline some of their weaknesses. With respect to the event-study approach, see also Thornton (2014b).
[74] Some argue that lower bond yields may as well be explained by a global downward trend in long-term interest rates (Belke, Gros and Osowski, 2016).

review of the existing empirical evidence with respect to the impact of US quantitative easing on longer-term government bond yields in the US and abroad.

In order to estimate the effects of unconventional monetary policy on various nominal longer-term yields, Wright (2012) applies the methodology developed by Rigobon and Sack (2004). The period of investigation includes the selected FOMC meetings and speeches during 2008:11 – 2011:9. The structural VAR model, containing daily financial data, is identified by assuming the heteroscedasticity of monetary policy shocks, i.e. the variance of monetary policy shocks on the days of policy meetings and certain speeches is particularly high as compared to the remaining days, while other structural shocks are assumed not to exhibit such heteroscedasticity. The monetary policy shock is normalised to lower the 10-year Treasury yield by 25 basis points immediately. The impulse responses of the 10-year Treasury, AAA and BAA corporate bond yields show that the long-term rates decline and this reaction is also statistically significant. Nevertheless, it is significant for only a short period of time as the impact dies out relatively quickly. Meanwhile, the 2-year yield decreases moderately and the decline is not statistically significant. Similar findings are obtained using a high-frequency event study over the same set of monetary policy events. The results show that a one-standard-deviation monetary policy shock reduces the 10- and 2-year Treasury yields by around 12 and 6 basis points, respectively (Wright, 2012). These effects on yields are highly statistically significant. Furthermore, the event-study analysis provides the evidence of the international effects of US unconventional monetary policies, i.e. yields on long-term government bonds in Canada, the UK and Germany also decline and this decline is statistically significant.

The international effects of the Fed's quantitative easing have been also documented by several other studies. Fratzscher, Lo Duca, and Straub (2013) investigate the impact of QE and the liquidity facilities on asset prices and capital flows in the US and 65 foreign countries. Using a daily panel regression framework, they distinguish between the QE-related announcements and actual Fed's intervention, i.e. liquidity provision and the purchases of longer-term Treasury bonds, agency debt and MBS. The model is estimated over the period 2007:1 – 2010:12. With respect to announcements, the first round of asset purchases resulted in a significant reduction in the 10-year government bond yields domestically and globally. On the contrary, QE2 had a much smaller, albeit still significant, impact on the 10-year Treasury yield; however, it did not have any impact on international bond yields. In addition, liquidity operations also were successful in reducing the 10-year Treasury yield and, to some smaller extent, yields in advanced economies. While MBS purchases had no significant effect, the purchases of Treasury securities raised slightly the US yield. Fratzscher, Lo Duca, and Straub (2013) suggest that these findings are in line with the portfolio balance channel. They show that QE1 triggered capital inflows into US bonds and equities and contributed to outflows out of emerging market economies. In contrast, QE2 caused capital to flow out of the US bond and equity markets into foreign equity markets, in addition to a general outflow from bonds into stocks within countries.[75]

Using the event-study approach, Neely (2015) demonstrates that the announcements related to the Fed's unconventional monetary policies in 2008-2009 significantly compressed nominal yields on long-term government bonds not only in the US but also in Australia, Canada, Germany, Japan and the UK. In the second part of the analysis, a simple portfolio balance model of bond returns is used to provide some insight as to whether the decline in yields can be explained through lower expected return on bonds, i.e. the term premium component. The estimated model predicts that central bank asset purchases should reduce expected excess bond returns. As the actual data is also consistent with these predictions, the portfolio balance channel is likely to explain the decline in yields associated with US unconventional monetary policy (Neely, 2015).

The event study by Gagnon et al. (2011) is focused on the effects of the selected QE-related announcements between November 2008 and March 2010 on the Treasury yields and the term premium, among other long-term interest rates. The cumulative yield changes around the baseline set of events show that the 2- and 10-year Treasury yields declined substantially and the response of the long-term yield was much stronger. The response of the 10-year term premium, which is estimated using the model of Kim and Wright (2005), is similar in magnitude to the overall decline in the long-term yield. This implies that the impact on the long-term yield is largely explained by the reduction in the term premium component, providing support for the portfolio balance channel (Gagnon et al., 2011). In addition, the pre-LSAP time series analysis of the historical developments in the term premium on the 10-year Treasury bond is provided. The term premium is modelled as a function of business cycle factors, the net public sector supply of debt securities with long maturities, and factors related to uncertainty surrounding economic fundamentals over the period 1985:1 – 2008:6. The estimated coefficients across alternative specifications suggest that $1.725 trillion assets purchased by the Fed in 2008 – 2010 could have reduced the 10-year term premium by about 38 to 82 basis points (Gagnon et al., 2011).

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[75] The analysis of stock returns indicates that the announcements related to QE1 and QE2 led to significantly higher US stock prices. In addition, QE2 increased stock prices in foreign markets.

Several studies take the advantage of the security-level data and analyse the portfolio balance channel in more detail, i.e. they distinguish between the local supply and duration effects. For instance, D'Amico et al. (2012) analyse the impact of the Treasury bond purchases associated with QE1 and QE2 programmes on medium- to long-term nominal Treasury yields. In order to measure the effects on the nominal term premium and its components, i.e. the real term premium and inflation risk premium, nominal yields are decomposed using several term structure models. The regression analysis of Treasury yields in the pre-LSAP period shows that their constructed measures of the aggregate duration and local supply of Treasuries are both important explanatory variables carrying positive and significant coefficients. From the term premium regressions it can be noted that the major part of the local supply and duration effects on nominal yields is transmitted via the term premium (nominal or real) component. Based on their preferred model specification, D'Amico et al. (2012) suggest that Treasury purchases reduced longer-term yields through both the duration and the local supply channels, in addition to any potential signalling channel effect.

D'Amico and King (2013) test for the local supply effect, i.e. the response of Treasury yields with a given maturity to changes in the supply of securities with that maturity, across the yield curve with respect to the Treasury purchases during QE1. Furthermore, the regression analysis of returns on Treasury securities also allows partitioning the local supply effect into the "stock" and "flow" effects.[76] The findings show that yields on specific securities declined following the purchases of these securities and the securities of similar maturity, supporting the local supply channel. Based on the estimations, D'Amico and King (2013) construct the counterfactual yield curves by deducting the impact of the stock effect from the actual Treasury prices as of the end of QE1. On average, Treasury yields declined by around 30 basis points over the course of the programme. The flow effect resulted in a further decline in the yields of purchased securities of around 3.5 basis points on the days of these purchases. Overall, this study provides support for the preferred-habitat theory in explaining the LSAPs effects on longer-term interest rates.

The study by Cahill et al. (2013), also using the security-level data, examines the relative importance of the duration and local supply channels in explaining the decline in long-term yields. They estimate the cross-section regressions of changes in yields for each of the selected announcements related to QE1, QE2 and MEP programmes and a pooled

[76] The stock effect refers to the aggregate effect on Treasury yields of all relevant asset purchase operations between 17 March 2009 and 30 October 2009, while the flow effect reflects immediate changes in yields at the time of purchases.

data regression for all announcements together. After fully accounting for the pre-announcement market expectations, the results indicate that the duration and local supply channels explain almost the entire variation in Treasury yields. Furthermore, both channels appear to be equally important with respect to yield changes.

With respect to the portfolio balance channel, Krogstrup, Reynard and Sutter (2012) argue that there may have been the liquidity effect at work, in addition to the local supply effect. This implies that changes in central bank liabilities, induced by the asset purchases that have been financed with central bank money, have an impact on bond yields.[77] Within a single framework, they analyse both the local supply and liquidity effects. The regressions of the 10- and 5-year Treasury yields are estimated with weekly data over the period 1990:2 – 2011:1. The estimation results show that changes in the supply of Treasuries available to the public and changes in bank reserves both contribute to a decline in yields. Overall, the liquidity effect may have reduced longer-term interest rates by around 46 - 85 basis points between January 2009 and January 2011 (Krogstrup, Reynard and Sutter, 2012).

On the other hand, as noted by Thornton (2012), the theoretical basis for the portfolio balance channel, typically emphasised in the literature, is somewhat weak and requires a strong assumption about investors' preferences. In the empirical analysis, he closely follows Gagnon et al. (2011) and examines the portfolio balance effect of quantitative easing for the US. In addition to the original variables, the slope of term structure is considered as an alternative dependent variable and a larger number of measures of government debt supply is used in the regressions. Thornton (2012) argues that the results based on the specification as in Gagnon et al. (2011) are driven by common trends in the term premium and the public debt supply. After including the trend in regressions, there is no evidence for the portfolio balance channel, i.e. the supply of debt securities held by public is not important in the yield or term premium equations. Thus, possibly, the signalling channel has been understated in the literature (Thornton, 2012).

In line with Thornton (2012), other studies also point out a potentially relevant contribution of the signalling effect of quantitative easing. For instance, Krishnamurthy and Vissing-Jorgensen (2011) argue in favour of the signalling channel, in addition to the

[77] This type of the liquidity channel is distinct from the market functioning channel discussed by Gagnon et al. (2011). According to the liquidity channel, banks are left with more reserves following asset purchases by a central bank and trade their excess reserves for other (positive yielding) assets. In turn, the prices of those assets increase and yields decrease. This channel therefore also relies upon the imperfect substitutability of assets, making it somewhat similar to the portfolio balance channel.

"safety" channel.[78] They examine the responses of Treasury bond and other longer-term yields to the selected announcements associated with QE1 and QE2 programmes. Regarding the first round of asset purchases, long-term yields decreased across the term structure with the response being stronger at the longer-end of the yield curve. The cumulative effects over the five selected events are found to be negative and typically significant. To gain some insight about the transmission channels, Krishnamurthy and Vissing-Jorgensen (2011) take a model-free approach. According to changes in the yields on federal funds rate futures around the event dates, the signalling channel is found to be very important in explaining the decline in Treasury yields. Furthermore, they separately evaluate the inflation channel using market-based measures of inflation expectations.[79] The QE announcements appear to be associated with an increase in inflation expectations and reduced inflation uncertainty. With respect to QE2, the effects on yields appear to be much smaller. Nevertheless, the evidence is still supportive of the signalling, safety and inflation channels.

Within the event-study framework, Christensen and Rudebusch (2012) evaluate the contribution of the portfolio balance and signalling channels to the decline in government bond yields following quantitative easing policies in the UK and US. For the US, eight announcements between November 2008 and November 2009 are selected. The total net reduction in the yields on 5- and 10-year Treasury bonds is 97 and 89 basis points, respectively. On the other hand, the 1- and 2-year yields declined by much less. Using four different dynamic term structure models, the response of yields is then decomposed into three components, i.e. changes in future expected short-term interest rates, the term premium and the residual term. Based on the preferred model, the decomposition of the 10-year yield response reveals that more than a half of the cumulative decline in the Treasury yield is accounted for by the decline in the expected future interest rate component and only about a third of the yield change is due to the reduced term premium (Christensen and Rudebusch, 2012). Thus, the signalling channel appears to be relatively strong in the US. On the other hand, the results for the UK show that the cumulative decline in the 10-year gilt yield over the QE announcements between February 2009 and October 2011 is largely explained by the decline in the term premium component, indicative of the strong portfolio balance effect (Christensen and Rudebusch, 2012).

[78] The safety channel refers to a part of the broader portfolio balance channel. It can be thought of as the portfolio balance channel for safe long-term assets (Krishnamurthy and Vissing-Jorgensen, 2011). As also discussed in D'Amico et al. (2012), the safety channel is subsumed within the scarcity (local supply) channel.

[79] Nevertheless, D'Amico et al. (2012) note that the inflation channel should not be considered in its own right since the response of inflation expectations is potentially the result of other channels.

Bauer and Rudebusch (2014) also consider both channels and evaluate the relative contribution of each with the focus on QE1 since further rounds of quantitative easing were largely anticipated by market participants. Specifically, the same eight announcements are considered as in the baseline analysis of Gagnon et al. (2011). Initially, they take a model-free approach and explain changes in Treasury yields around the announcement dates based on the money market futures rates and longer-term overnight index swap rates. Both types of data indicate the signalling channel since the announcements are associated with lower expectations about future short-term rates. In the second part of the analysis, changes in the yields on 5- and 10-year Treasury bonds are decomposed into changes in short-term interest rate expectations, i.e. the expected policy rate, and changes in the term premium using a variety of term structure models. As discussed by Bauer and Rudebusch (2014), standard dynamic structure models suffer from several statistical problems. In this study, they address two key issues, i.e. the small-sample bias and statistical uncertainty. Consequently, the results imply that the role of the signalling channel is much greater than is typically reported in other studies. The actual contribution of the signalling channel in explaining yield changes is likely to be around 40% - 50%, as opposed to approximately 22% based on the estimates in Gagnon et al. (2011). This is consistent with the view that, through the announcement and implementation of longer-term asset purchases, the Fed has signalled to the market that it would conduct expansionary monetary policy for longer than previously expected (Bauer and Rudebusch, 2014).

Bauer and Neely (2014) extend the analysis of US quantitative easing transmission to international bond markets. With respect to all three rounds of quantitative easing, they evaluate the contribution of the signalling and portfolio balance channels to the observed reduction in domestic and foreign government bond yields. The results from an event study show that QE1 announcements are associated with significant declines in the 2- and 10-year government bond yields across the countries with the largest effects felt in the US and Canada. In contrast, the announcements related to QE2 and QE3 had much smaller and typically insignificant impact on the US and international yields. The model-free analysis of OIS rates implies that QE-related announcements reduced long-term yields to a large extent through decreases in future expected policy rates. Using a variety of dynamic term structure models, the changes in the 10-years yields on the announcement days are decomposed into changes in short-term rate expectations and the term premium for each country. With respect to all rounds of asset purchases, the signalling channel explains between 45% and 90% of the total decline in the 10-year Treasury yield. Similarly, very strong signalling effects are found for the 10-year government yield in Canada. In the case of Australia and Germany, the contribution of expectations about a short-term rate to the decline in long-term yields is smaller but still significant, while it is negligible for Japan.

Wu (2014) further strengthens the argument in favour of the signalling channel. The nominal 10-year Treasury yield is decomposed using the model by Kim and Wright (2005) and a time series regression is estimated for each component of the long-term yield over the period 1992:1 – 2013:9. The explanatory variables include macroeconomic and financial variables as well as the constructed real-time measure of LSAP-related policy, which denotes market expectations about the size and persistence of asset purchases. The results show that unconventional monetary policies reduced significantly the term premium on the 10-Treasury yield as well as lowered the expected path of future short-term interest rates. The portfolio balance and signalling channels both appear to be active. Also, there is the evidence of spillover effects between the two channels. The Fed's forward guidance contributes to lower term premium as it leads to the gradual extension of market expectations about holding period of purchased assets. On the other hand, the LSAPs help reduce the expectations component of yields, potentially due to the increased credibility of the forward guidance (Wu, 2014). In line with the previous studies, the findings also indicate that the final round of quantitative easing was the least effective in reducing longer-term yields.

In general, it is found that quantitative easing policy implemented by the Fed significantly lowered longer-term yields on government bonds in the US as well as bond yields in foreign countries. Nevertheless, the impact of the second and third rounds of the LSAPs appears to be much smaller and less persistent. Finally, the literature has not yet reached the consensus on which channel, the portfolio balance or signalling, explain the movements in Treasury yields in response to unconventional monetary policy shocks. It is also likely that both channels are at work and may be equally important in explaining changes in longer-term interest rates induced by the central bank's asset purchase programmes.

Several points should be noted. Firstly, the literature that applies the VAR-based returns variance decomposition methodology to examine monetary policy effects on the bond market is rather scant. Secondly, the existing studies are focused on long-term bonds. Thirdly, the sample period examined in these studies typically starts in the 1970s and ends prior to the global financial crisis. Thus, the impact of the crisis itself and the effects of unconventional monetary policy tools are not considered within this framework. Finally, there is still a debate about which channels of transmission explain quantitative easing effects on bond yields. To this respect, the methodology applied in Bredin, Hyde and

O'Reilly (2010) allows tracing down the component that is driving the response of bond returns to conventional monetary policy shocks. Therefore, it may be useful to gain some insight to the channels of quantitative easing effects on bond returns using this approach. The analysis here attempts to fill this gap. Using a similar approach to Bredin, Hyde and O'Reilly (2010), it firstly provides the variance decomposition of unexpected excess returns on medium- and long-term US government bonds. Secondly, it investigates the impact of US conventional and unconventional monetary policy on bond returns and their components.

2.5 Methodology

2.5.1 Excess bond returns decomposition

Using the framework of Campbell and Ammer (1993), current period unexpected excess bond returns are decomposed into the sums of revisions in expectations about future one-period excess bond returns (x), inflation (π) and real interest rates (r^i):

$$\tilde{x}_{n,t+1} = (E_{t+1} - E_t)\left[-\sum_{j=1}^{n-1} x_{n-j,t+1+j} - \sum_{j=1}^{n-1} \pi_{t+1+j} - \sum_{j=1}^{n-1} r^i_{t+1+j}\right] = -\tilde{x}_{x,t+1} - \tilde{x}_{\pi,t+1} - \tilde{x}_{r^i,t+1} \tag{2.1}$$

where $\tilde{x}_{n,t+1} = x_{n,t+1} - E_t\left[x_{n,t+1}\right]$ represents the unexpected one-period log return on a n-period zero-coupon bond in excess of continuously compounded one-period nominal interest rate (the bond becomes (n – 1)-period bond at t + 1), $\tilde{x}_{x,t+1}$ denotes revisions in expectations regarding future excess bond returns (*risk premium news*)[80], $\tilde{x}_{\pi,t+1}$ represents revisions in expectations about future inflation (*inflation news*) and $\tilde{x}_{r^i,t+1}$ denotes revisions in expectations regarding future real interest rates (*real interest rate news*).[81]

The decomposition implies that positive unexpected excess bond returns must be associated with decreases in expected future excess returns during the life of the bond, decreases in expected future inflation rates, decreases in expected future real interest rates, or the combination of the three. Equation (2.1) is a dynamic accounting identity that arises

[80] In line with the related literature, news about future excess bond returns and risk (term) premium news are used in the text interchangibly (see Campbell and Ammer, 1993; Barr and Pesaran, 1997; Engsted and Tanggaard, 2007).

[81] E_t represents expectations formed at the end of period t. See Appendix A for the derivation of this identity.

from the definition of bond returns and imposes internal consistency on expectations.[82] It is not a behavioural model containing economic theory and asset pricing assumptions. Nevertheless, both the Fisher hypothesis and the expectations theory of the term structure have important implications for the decomposition of excess bond returns. Specifically, the former hypothesis implies that *ex ante* real interest rates are constant and therefore the real interest rate news term is zero. The latter hypothesis assumes time-invariant expected excess bond returns which are consistent with the risk premium news term being zero. Therefore, in the extreme, if both hypotheses hold, inflation news will be the only source of the variation in bond returns in excess of a short-term risk-free rate.[83]

From Equation (2.1) it follows that the total variance of unexpected excess returns can be decomposed into the sum of the three variance terms plus the respective covariance terms between the components:

$$
\begin{aligned}
Var\left(\tilde{x}_{n,t+1}\right) &= Var\left(\tilde{x}_{x,t+1}\right) + Var\left(\tilde{x}_{\pi,t+1}\right) + Var\left(\tilde{x}_{r^i,t+1}\right) + 2Cov\left(\tilde{x}_{x,t+1}, \tilde{x}_{\pi,t+1}\right) \\
&+2Cov\left(\tilde{x}_{x,t+1}, \tilde{x}_{r^i,t+1}\right) + 2Cov\left(\tilde{x}_{r^i,t+1}, \tilde{x}_{\pi,t+1}\right)
\end{aligned}
\tag{2.2}
$$

In order to evaluate the relative importance of news about future excess bond returns, inflation and real interest rates, each variance and covariance term in Equation (2.2) is normalised by the total variability of excess returns.

2.5.2 Vector autoregressive model and news

The implementation of the variance decomposition for excess bond returns requires empirical proxies for directly unobservable revisions in expectations regarding future excess returns, inflation and real interest rates. Campbell and Ammer's (1993) methodology links these multi-period expectations to the stationary dynamics of a vector autoregressive model. Specifically, a first-order reduced-form VAR is employed, involving the variables of interest along with other indicators that may be useful in forecasting them,

[82] Unlike in the case of stocks, the dynamic accounting identity for zero-coupon bonds holds exactly rather than approximation.

[83] Existing evidence regarding the empirical validity of the expectations hypothesis and the Fisher hypothesis can be described as mixed with the role of the adopted testing procedures being crucial. Sarno, Thornton and Valente (2007) use a more powerful test with either macroeconomic factors or more than two bond yields and overturn evidence from conventional tests by showing that the expectations hypothesis can be rejected throughout the maturity spectrum. Christopoulos and Leon-Ledesma (2007) attribute the lack of widespread empirical evidence for the Fisher hypothesis in cointegration-based studies to non-linearities in the long-run relationship between nominal interest rates and inflation.

to obtain empirical proxies for the news components in Equation (2.1).[84] The forecast errors and the estimated parameters from the VAR model are used to construct the time series of revisions in expectations for the variables of interest. The calculation of these empirical proxies does not depend on the ordering of the state variables since it is based on the reduced-form residuals.

With respect to the standard errors for the terms of the variance decomposition, this chapter follows the approach in Campbell and Ammer (1993).[85] The VAR coefficients and the elements of the variance-covariance matrix of residuals are jointly estimated by the generalised method of moments. The coefficient estimates are identical to the standard OLS estimates. However, this method delivers the heteroscedasticity-consistent variance-covariance matrix (V) for the full set of the VAR parameters (γ). The terms of the variance decomposition in Equation (2.2) are nonlinear functions of the estimated VAR parameters, i.e. $f(\gamma)$. Then, the standard errors for these variance terms can be obtained utilising the delta method and computed as $\sqrt{f_{\gamma}(\gamma)' V f_{\gamma}(\gamma)}$.[86]

The starting point is the definition of a state vector containing stationary variables that help to measure and forecast excess bond returns, inflation and real interest rates:

$$Z_{t+1} = AZ_t + W_{t+1} \tag{2.3}$$

where Z_t is the state vector of endogenous variables included in the model, A denotes the matrix of VAR parameters, and W_{t+1} is the vector of forecast residuals. The state vector includes the first difference of the nominal short-term risk-free rate ($\Delta y_{1,t}$), the spread between long-term and short-term yields ($s_{n,t}$), the real interest rate (r_t^i), the relative bill rate (rb_t), i.e. the difference between the nominal short-term interest rate and its 12-month backwards moving average.[87]

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[84] The VAR(1) assumption is not restrictive. The robustness analysis shows that the findings obtained using the VAR(1) model are robust to the use of higher order VARs.

[85] This approach is also employed by Barr and Pesaran (1997) and Bernanke and Kuttner (2005).

[86] This chapter uses the RATS code for the replication of Campbell and Ammer (1993) kindly made available by Tom Doan at RATS online forum. With respect to the matrix of partial derivatives, it is derived using numerical derivatives.

[87] Three separate VAR models are estimated for three maturities as a "long-term" bond in the term spread: 2-, 5- and 10-year zero coupon bonds. This is related to a parsimonious 3-factor yield curve model by Nelson and Siegel (1987). A flexible, smooth parametric function provides an approximation of the zero-coupon yield curve describing the relationship among short-, medium, and long-term yields. The three factors in the model, i.e. the long-, short- and medium-term components of the yield curve, capture well the shapes of the term structure typically observed in data over time (Nelson and Siegel, 1987; Moench, 2012). Diebold and Li (2006) interprets these factors as dynamic latent level, slope and curvature factors. The loading on the long-term factor is a constant and is the same at all maturites. Thus, an increase in this factor affects all yields equally, consequently changing the level of the yield curve. The short-term factor is also known as the slope factor. Its loading decays to zero as maturity increases implying that the loading is greater on short-term rates. Hence, an increase in this factor affects short-term rates more than long-term rates, thereby changing

The first two variables in the state vector are used to construct innovations in excess bond returns. The term spread has strong predictive power over bond returns (Campbell and Shiller, 1991; Fama and Bliss, 1987; Greenwood and Vayanos, 2014), while the relative bill rate is a forecasting variable that can capture longer-run dynamics of interest rate changes without introducing long lags (Campbell and Ammer, 1993; Barr and Pesaran, 1997; Bernanke and Kuttner, 2005). The estimated VAR allows computing unexpected excess bond returns and the three components in Equation (2.1) as follows:

$$\tilde{x}_{n,t+1} = -(n-1)(s_1^T W_{t+1} + s_2^T W_{t+1}) \,, \tag{2.4}$$

$$\tilde{x}_{r^i,t+1} = s_3^T (I-A)^{-1}(A-A^n)W_{t+1} \,, \tag{2.5}$$

$$\tilde{x}_{\pi,t+1} = -\tilde{x}_{r^i,t+1} + s_1^T \left\{ \left(I-A\right)^{-1} \left[\left(n-1\right) I + \left(I-A\right)^{-1} \left(A^n - A\right) \right] \right\} W_{t+1} \,, \tag{2.6}$$

$$\tilde{x}_{x,t+1} = -\tilde{x}_{n,t+1} - \tilde{x}_{r^i,t+1} - \tilde{x}_{\pi,t+1} \,. \tag{2.7}$$

where s_i^T is the unit vector with i representing the i^{th} equation in the model and accordingly the i^{th} element of the vector is set to 1; I is the identity matrix.[88]

Equation (2.4) shows that current unexpected excess bond returns are obtained using innovations from two VAR equations: the change of nominal short-term rate and the term spread. The inclusion of the real interest rate in the state vector allows the extraction of news about it directly from the model as indicated by Equation (2.5). In Equation (2.6), the inflation news term is computed by combining innovations in the change of the nominal short-term rate with real interest rate news. Finally, Equation (2.7) shows that risk premium news, i.e. the sum of revisions in expectations about future excess bond returns, is obtained as the residual term using the dynamic accounting identity and the estimates of other components. This is necessary for *n*-period zero-coupon bonds since shrinking maturity over the life of the bond precludes the direct forecasting of excess returns using the VAR model. Hence, excess bond returns are not directly included in the state vector and the related news component is backed out as the residual term. Alternatively, one may take a slightly different approach by calculating excess bond returns using a bond index and including returns directly into the VAR. This way, inflation news is backed out as a

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the slope of the term structure. Finally, the medium-term factor loads more heavily on the medium-term yields and less on the short- and long-term rates. An increase in the third factor will increase the medium-term yields but will only have a small effect on other rates, i.e. the curvature of the term structure changes.

[88] See Appendix B for more details.

residual component instead (Bredin, Hyde and O'Reilly, 2010). Nevertheless, this would require assuming the perpetuity of 10-year government bonds.

As Engsted, Pedersen and Tanggaard (2012) explain, the need to account for the shrinking maturity is crucial within the framework adopted in this chapter. Ignoring this may lead to unwarranted conclusions about the reliability of the bond market variance decomposition as in Chen and Zhao (2009). Chen and Zhao (2009) argue that since the nominal cash flows of Treasury bonds are fixed, the estimated cash flow news (of which inflation news is a component) must be zero. Thus, news about future expected excess returns should be driving current unexpected returns. To illustrate, they decompose unexpected excess bond returns on Treasuries into two components: cash flow news and risk premium news, where the former is backed out as the residual from the VAR estimation. The findings are inconsistent with the prediction as the estimated variance of cash flows news is not zero, and is at least as large as that of risk premium news. Chen and Zhao (2009) attribute this to the problem of omitted relevant state variables to forecast excess returns. However, as Engsted, Pedersen and Tanggaard (2012) point out, Chen and Zhao (2009) neglect the shrinking maturity of the bonds over their lifetime. Furthermore, while they analyse excess bond returns in the VAR, the formula that they use for the decomposition holds for raw returns only.

The VAR model that is estimated to obtain the news components is assumed to contain all relevant information that investors may have when forming expectations about the future. Given variability in the components of excess bond returns, the variance decomposition is indeed conditional upon this information. If investors have additional information that is not present in the state vector, the relative importance of the residual component (news about future excess bond returns in this analysis) may be overstated. In the robustness analysis section it is demonstrated that the baseline findings, based on the state vector described above, are robust to the incorporation of additional macro-financial predictor variables in the state vector. Furthermore, as it is discussed later, the key component that drives bond returns variability is not the residual term. Thus, it is less likely that the results are driven by the misspecification of the model.

2.5.3 Monetary policy effects

The above sections explain how the variance of unexpected excess bond returns can be linked to news about future excess returns, inflation and real interest rates, and how these news terms can be obtained from the estimated VAR model. This section presents the framework that is used to estimate the impact of monetary policy actions on the bond market. To do so, the extension of Campbell and Ammer's (1993) methodology by Bernanke and Kuttner (2005) is modified for the case of the bond market. [89] This approach generates the estimates of the impact of monetary policy actions on unexpected excess bond returns and the related news components, thereby providing insights to the sources of the bond market's response to monetary policy. The starting point is the inclusion of a monetary policy indicator (*MP*) as an exogenous variable in the VAR(1) model:

$$Z_{t+1} = AZ_t + \phi MP_{t+1} + W^*_{t+1} \tag{2.8}$$

where ϕ is the vector of the response parameters of state variables to contemporaneous monetary policy actions. As it is explained in Section 2.6.3, four alternative monetary policy indicators are employed that relate to actual and surprise changes in the policy rate and the quantity of money.

From the above it follows that the original VAR forecast error vector W_{t+1} in Equation (2.3) is decomposed into the component related to monetary policy actions (ϕMP_{t+1}), and the component related to other information (W^*_{t+1}). Firstly, the original VAR(1) model is estimated to obtain the estimates of *A*. Secondly, the one-step-ahead forecast residuals are regressed on the monetary policy indicator variable in order to estimate ϕ. This two-stage procedure is preferred over the direct estimation of Equation (2.8) as it allows estimating the VAR dynamics over the longer sample period than the period used for monetary policy regressions, increasing the precision.[90]

It is now possible to calculate the effect of monetary policy on current unexpected excess returns and on news about expected real interest rates, inflation and excess bond returns, i.e. the risk (term) premium. In Equations (2.4)-(2.6), W_{t+1} is replaced with $\phi MP_{t+1} + W^*_{t+1}$ and partial derivatives with respect to MP_{t+1} are computed. Then, the responses of bond returns and three components to monetary policy shifts can be written as follows:

$$\tilde{x}^{MP}_{n,t+1} = -(n-1)(s_1^T + s_2^T)\phi \tag{2.9}$$

[89] As mentioned earlier, Bredin, Hyde and O'Reilly (2010) also consider the impact of monetary policy on bond returns and their components using Bernanke and Kuttner's (2005) VAR-based approach. Their analytical framework, however, is different. The formulas used for the decompositions of bond returns apply to the case of infinite maturity coupon bonds. This allows them to include excess returns directly in the VAR.

[90] The monetary policy effects are estimated for an alternative (shorter) sample period in the robustness analysis.

$$\tilde{x}^{MP}_{r^i,t+1} = s_3^T (I - A)^{-1} (A - A^n)\phi \tag{2.10}$$

$$\tilde{x}^{MP}_{\pi,t+1} = -s_3^T (I - A)^{-1} (A - A^n)\phi + s_1^T \left\{ (I - A)^{-1} \left[(n-1) I + (I - A)^{-1} \left(A^n - A \right) \right] \right\} \phi \tag{2.11}$$

$$\tilde{x}^{MP}_{x,t+1} = -\tilde{x}^{MP}_{n,t+1} - \tilde{x}^{MP}_{\pi,t+1} - \tilde{x}^{MP}_{r^i,t+1} \tag{2.12}$$

As in Bernanke and Kutner (2005), the delta method is used to compute standard errors for these responses.

2.6 Data and sample period

2.6.1 Sample period

The empirical analysis is based on monthly US data over the period 1985:1 – 2014:2. The sample commences during the early years of the Great Moderation period, while its latter part contains the recent global financial crisis and its aftermath. The estimations are conducted over both the full sample period (1985:1 – 2014:2) and a shorter sample (1985:1 – 2007:7) that ends prior to the onset of the recent financial crisis.[91] Doing so, allows getting some insight about the impact of the crisis on the variance decomposition of unexpected excess bond returns and the relationship between monetary policy actions and bond returns.

2.6.2 VAR state variables

The 1-month Treasury bill rate, obtained from the Centre for Research in Security Prices (CRSP), is used as a proxy for the nominal short-term risk-free interest rate ($y_{1,t}$). The long-short spreads ($s_{n,t}$) are calculated as the difference between the 10-, 5-, and 2-year zero-coupon Treasury bond yields and $y_{1,t}$. Data on continuously compounded zero-coupon yields is obtained from the daily dataset provided by Gurkaynak, Sack, and Wright (2007).[92] The *ex post* real interest rate is defined as the difference between $y_{1,t\text{-}1}$ and the current monthly inflation rate, measured by the change in the log of seasonally adjusted Consumer Price Index All items (CPI). The CPI data is provided by the Federal Reserve

[91] The start of the financial crisis is dated to August 2007 when doubts about global financial stability emerged and the first major central bank interventions in response to increasing interbank market pressures took place (Brunnermeier, 2009; Kontonikas, MacDonald and Saggu, 2013).
[92] The dataset is available online at http://www.federalreserve.gov/pubs/feds/2006/200628/200628abs.html.

Bank of St Louis (FRED database). The relative bill rate is calculated as the deviation of $y_{1,t}$ from its 12-month backwards moving average. All state variables are expressed in percentages per annum on continuously compounded basis (end of month data is used).[93]

2.6.3 Monetary policy indicators

Both the Fed's operating procedures and underlying macro-financial environment have changed over time. By the early 1980s, Volcker's disinflation was largely accomplished with inflation sharply reduced to around 3% in 1983. This development allowed interest rates to decline and eventually ushered the Great Moderation era that was characterised by overall macroeconomic stability. Monetary policy conduct during that period was characterised by the federal funds rate targeting and increasing transparency, with the Fed announcing its decisions for the target FFR after each FOMC meeting since February 1994.[94] The financial crisis of 2007-2009 brought this benign regime to an end and had a significant impact on the Fed's approach to monetary policy implementation. The Fed responded aggressively to the crisis by reducing the target FFR to near zero. Moreover, it used various tools (liquidity facilities and quantitative easing) to improve financial market conditions and put downward pressure on longer-term interest rates, thereby supporting economic activity.[95]

Conducting the LSAPs programmes, the Fed purchased significant amounts of longer-term assets from the private sector, mainly Treasury bonds and agency mortgage backed securities, leading to significant changes in the size and the composition of its balance sheet. The holdings of short-term Treasuries have declined due to the initial sterilisation of liquidity operations and the Maturity Extension Programme that followed later on. Meanwhile, longer-term securities held outright have significantly increased reflecting changes in the nature and the scope of the Fed's Open Market Operations (OMOs) as a result of the LSAPs.[96] The increase in the Fed's assets was matched by the

[93] Figure C2.1 in Appendix C presents the state variables.

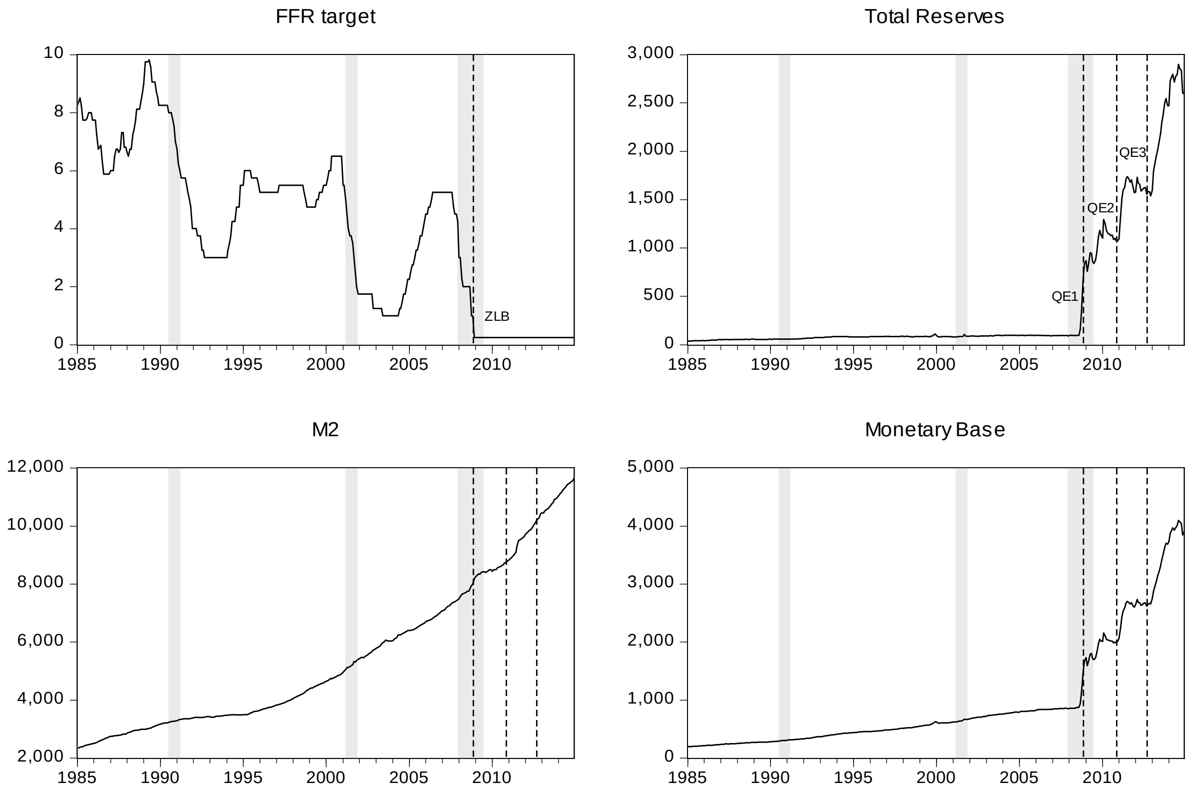
[94] US monetary policy operating procedures have included the period of targeting the FFR, i.e. the interest rate on overnight loans of reserves between banks, (1972–79 and 1988–present), non-borrowed reserves targeting (1979–82) and borrowed reserves targeting (1982–88). There is substantial empirical evidence indicating that the FFR is the key US monetary policy indicator during both the pre-1979 and post-1982 periods (Bernanke and Blinder, 1992; Bernanke and Mihov, 1998; Romer and Romer, 2004).

[95] These included (i) the provision of short-term term liquidity to banks and other financial institutions through the discount window lending and other facilities, such as the Term Auction Facility; (ii) the direct provision of liquidity to borrowers and investors in important credit markets via e.g. the Commercial Paper Funding Facility; (iii) the large-scale asset purchases that aimed to support credit markets and improve overall financial conditions. See Table C2.2 in Appendix C for the list of relevant announcements by the Fed.

[96] Figure C2.2 in Appendix C shows developments in the Fed's holdings of Treasury securities across different maturities. Note that, traditionally, OMOs involved the repurchase (repo) and sale-repurchase

expansion in its liabilities. Particularly, reserve balances have increased considerably relative to their level prior to the financial crisis and are highly in excess of the regulatory requirements. Reserves became the main component of the monetary base since currency in circulation continued to exhibit only a gradual increase over time. Figure 2.1 shows the dramatic rise in the total reserves and monetary base since late 2008. It also highlights that, in contrast to narrow money, broad money (M2) did not expand significantly. The lack of the more dramatic shift in broader monetary aggregates is related to the fact that banks let their excess reserves to increase sharply (Fawley and Neely, 2013).[97]

Figure 2.1: Policy rate and monetary aggregates



Notes: This figure plots the target federal funds rate (FFR target), the St. Louis adjusted total reserves (in $bn), the M2 money stock (in $bn) and the St. Louis adjusted monetary base (in $bn) over the full sample period (1985:1 – 2014:2). The dashed vertical line in the upper left panel denotes the start of the zero lower bound period. In the rest of the panels, the three dashed vertical lines denote the announcements of the first round of quantitative easing (QE1, 2008:11), the second round (QE2, 2010:11) and the third round (QE3, 2012:9). Shaded areas denote US recessions as classified by NBER business cycle dates. Data is obtained from FRED database.

(reverse repo) of securities, mainly short-term Treasuries, by the Fed in order to keep the FFR close to the target. Fama's (2013) empirical evidence indicates that indeed the FFR adjusts quickly towards the target.

[97] Fama (2013) attributes this development to the payment of interest on excess reserves by the Fed since October 2008, which implies that they no longer impose a cost on banks.

These developments renewed the focus of central bankers and monetary economists to quantity-based policy indicators with a number of recent theoretical (Curdia and Woodford, 2011; Gertler and Karadi, 2013) and empirical studies (Krogstrup, Reynard and Sutter, 2012; Gambacorta, Hofmann and Peersman, 2014) investigating the financial and macroeconomic role of quantitative easing and evaluating the monetary base, or the supply of reserves, as an alternative operating target.

The empirical analysis in this chapter uses four monetary policy indicators that are related to actual and unexpected changes in the federal funds rate and the (log) monetary base. The interest rate-based measures are capturing conventional monetary policy, while unconventional policy actions are gauged by the quantity-based measures.[98] The first indicator is changes in the FFR defined as $\Delta FFR_t = FFR_t - FFR_{t\text{-}1}$. This proxy is frequently utilised in several previous studies (Chen, 2007; Kontonikas and Kostakis, 2013; Maio, 2014).

The second indicator isolates unexpected (surprise) FFR changes using the data from the federal funds futures and the methodology developed by Kuttner (2001). The earlier literature that employs this proxy includes Bernanke and Kuttner (2005) and Bredin, Hyde and O'Reilly (2010), among others. The month-t unexpected FFR change (ΔFFR_t^U), is calculated as follows:

$$\Delta FFR_t^U = \frac{1}{D}\sum_{d=1}^{D} i_{t,d} - f_{t-1,D}^1 \tag{2.13}$$

where $i_{t,d}$ denotes the target federal funds rate on a day d of a month t, and $f_{t-1,D}^1$ is the rate corresponding to the 1-month federal funds futures contract on the last (D^{th}) day of the month t-1.[99]

The rate on the futures contract is defined as 100 minus the settlement price of the contract that is based on the average of the relevant month's effective FFR. Hence, the futures rate is a natural, market-based measure of expectations about future Fed's policy actions.[100] Gurkaynak, Sack and Swanson (2007) demonstrate that futures rates are

[98] After reaching the ZLB, the Fed has relied increasingly on the forward guidance. It is not an easy task to take into account the role of such policy tool separately given that monthly data is employed. Thus, note that the policy indicators used here does not account for the role of the forward guidance explicitly.

[99] The federal funds target rate was defined as an interval 0-0.25% in December 2008. As a result, the effective federal funds rate is used instead of the FFR target to denote $i_{t,d}$ between then and the end of the sample. Alternatively, the mid-point of the interval may be used, albeit the results do not change much. At the ZLB, given the Fed's forward guidance, monetary surprises should be close to zero.

[100] It should be noted that measuring surprise changes using an average FFR may understate the magnitude of policy surprises. The time-aggregation issue is analysed in Evans and Kuttner (1998).

dominant among financial market instruments with respect to forecasting the changes in the federal funds rate at horizons within six months.[101]

The third indicator is the growth rate of narrow money, measured by the change in the log of seasonally adjusted (St. Louis adjusted) monetary base (MB), $\Delta MB_t = MB_t - MB_{t-1}$. A number of studies that focus on the Japanese QE experience use developments in narrow money as a proxy for unconventional monetary policy (Kimura et al., 2003; Harada and Masujima, 2009). Monetary base developments should be more informative, as compared to asset-side measures, about the Fed's unconventional policies. This is because the asset side of the Fed's balance sheet just reflects LSAPs and shows a significant activity only since early 2009, while monetary base changes further capture the impact of non-sterilised liquidity provision through various facilities of the Fed that were heavily used in autumn 2008. Indeed, the highest monetary base growth rates occurred in October and November 2008 reaching 20% and 26% per month, respectively.[102]

The fourth indicator is based upon the previous work by Cover (1992) and Karras (2013). The surprises in narrow money growth (ΔMB_t^U) are obtained as the residuals from the regression of monetary base growth on its own lags and the lags of unemployment measure:

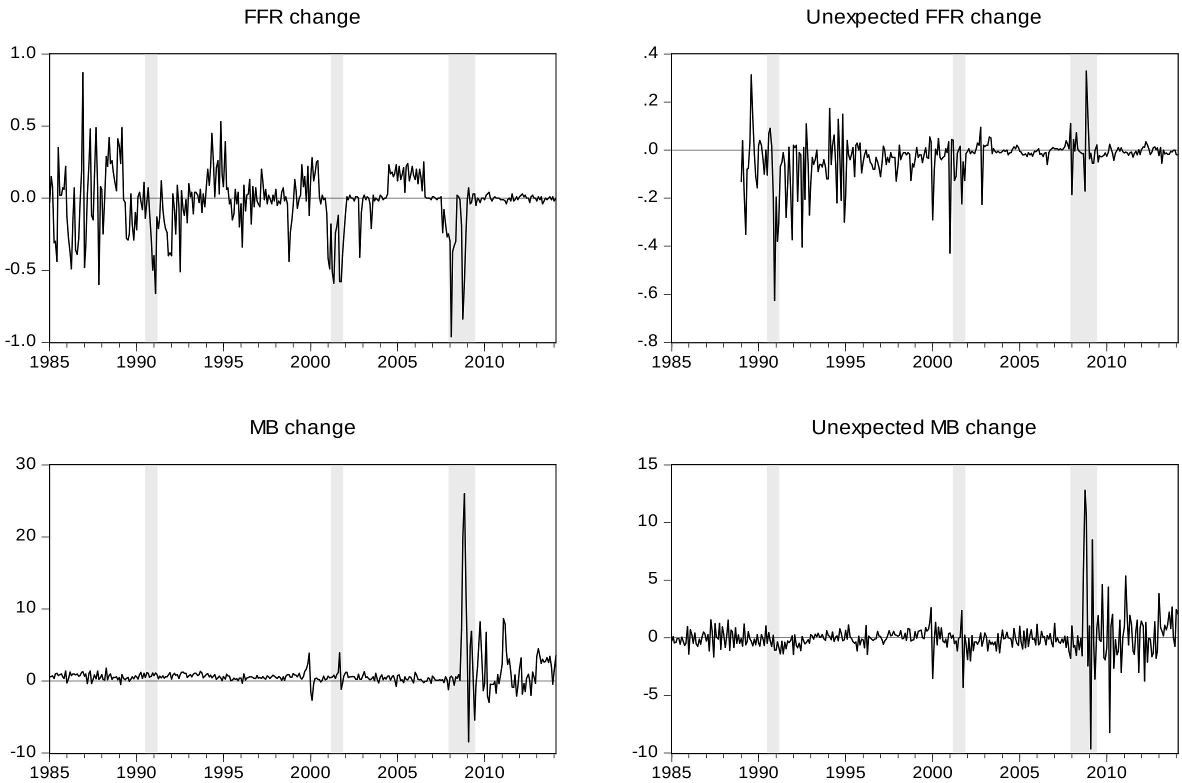
$$\Delta MB_t = a + \sum_{j=1}^{n} \beta_j \Delta MB_{t-j} + \sum_{i=1}^{m} \gamma_i UN_{t-i} + \varepsilon_t \tag{2.14}$$

where $UN_t = \log[U_t / (1 - U_t)]$ and U_t denotes the unemployment rate.[103]

The data for the FFR, monetary base and unemployment rate is obtained from FRED database, while federal funds futures rates are sourced from Bloomberg. Figure 2.2 plots all four monetary policy indicators. Towards the end of 2008, the quantity-based proxies become highly active, while the volatility of interest rate-based proxies displays a negative trend over time and dies out since the zero lower bound was reached.

[101] Note that federal funds futures contracts may carry the risk premium implying that the measures of monetary policy surprises using futures rates may be biased (Piazzesi and Swanson, 2008). Nevertheless, Cenesizoglu and Essid (2012) show that after adjusting monetary policy surprises for the potential risk premium, the estimated monetary policy effects on credit spreads do not change substantially.
[102] The corresponding figures for the total reserves growth were 78% and 66%. They also constitute historical highs.
[103] The number of lags (n=m=7) is chosen by the Akaike Information Criterion. The least squares estimation of Equation (2.14) indicates that monetary base growth is mainly explained by its own lags, with R^2 being equal to around 50%. In the robustness analysis section, several alternative empirical specifications for the monetary base growth are used but this does not change the baseline results.

Figure 2.2: Monetary policy indicators



Notes: This figure plots four indicators of monetary policy actions over the full sample period (1985:1 – 2014:2); the change in the Federal funds rate (FFR), the unexpected FFR change, the change in log monetary base (MB change) and the unexpected change in log monetary base. For further details, see Section 2.6.3. Shaded areas denote US recessions as classified by NBER business cycle dates.

2.6.4 Exogeneity assumption for monetary policy indicators

The indicator of monetary policy actions is included as an exogenous variable in Equation (2.8). However, the exogeneity assumption would not hold in the following three cases. First, if the Fed responds contemporaneously to developments in the market for Treasuries. Second, if the Fed and the Treasuries market jointly and contemporaneously respond to new information. Third, if policy actions reveal some private information that the Fed possesses about future economic developments, related to the superior resources that it commits to forecasting (Romer and Romer, 2000).[104] Previous studies have attempted to directly address the potential endogeneity problem in the relationship between monetary policy and asset prices by employing various empirical approaches.[105]

[104] For example, if expansionary monetary policy signals weaker economic outlook, commercial forecasters may respond by revising their inflation expectations downwards based on the inferred information from the Fed's actions, leading to lower yields and higher returns for bonds.

[105] One approach advocates the use of high-frequency data and the measurement of monetary policy shocks and market returns over a narrow time window around policy announcements. Thornton (2014a), however,

Nevertheless, as it is argued below, the exogeneity assumption should not be too restrictive, i.e. the assumption holds. In other words, the potential endogeneity issue with respect to the measure of monetary policy actions is not likely to have significant implications for the results reported here.

With respect to the first potential source of endogeneity, some argue that the Fed may be reacting to bond market developments due to forward-looking information about current and expected economic conditions likely to be reflected in bond prices (Piazzesi, 2005; Farka and DaSilva, 2011). Nevertheless, the empirical evidence on whether the Fed is systematically following Treasuries is overall inconclusive and rather elusive when medium- and longer-term yields, as the data used in this chapter, are examined (Nimark, 2008; Vazquez, Maria-Dolores and Londono, 2013). Second, in order to examine whether the policy indicators react to economic news, they are regressed on variables that capture surprises in nonfarm payrolls, industrial production growth, retail sales growth, core and headline CPI inflation (Bernanke and Kuttner, 2005). The results do not indicate any significant contemporaneous monetary policy response to macroeconomic surprises.[106] Furthermore, this analysis is focused on unexpected excess returns, thus, the direct monetary policy response to this measure is rather unlikely. Finally, the arguments made by Romer and Romer (2000) have been questioned. Faust, Swanson and Wright (2004) find little evidence that Fed policy surprises signal additional information about the state of economy or have any significant influence on the private sector forecasts.[107] Furthermore, Barakchian and Crowe (2013) demonstrate that even if monetary policy surprises are contaminated with the Fed's private information, the resulting simultaneity bias is likely to be small (see also Gertler and Karadi, 2015).

2.7 Empirical results

2.7.1 VAR estimation results

points out that using intraday data, as in Gurkaynak, Sack and Swanson (2005), the response of the market may reflect initial overreaction to monetary policy shifts. Instead, he proposes an approach based on daily data that helps to correct for the potential bias due to the joint response of monetary policy and the bond market to non-policy news. Alternatively, Rigobon and Sack (2004) suggest an approach based on the heteroskedasticity in high-frequency data associated with monetary policy actions.

[106] Due to data availability, the sample period for these regressions starts only in 1991:10. See Table C2.3 in Appendix C for the results.

[107] See also Faust and Wright (2008).

Table 2.1: VAR estimates

	10-year bonds					5-year bonds					2-year bonds				
	1985:1 – 2014:2														
	$\Delta y_{1,t}$	$s_{n,t}$	r_t^i	rb_t	R^2	$\Delta y_{1,t}$	$s_{n,t}$	r_t^i	rb_t	R^2	$\Delta y_{1,t}$	$s_{n,t}$	r_t^i	rb_t	R^2
$\Delta y_{1,t+1}$	-0.425***	0.085***	-0.010	0.103**	0.196	-0.414***	0.148***	-0.011	0.110***	0.236	-0.359***	0.271***	-0.017***	0.067*	0.322
	(0.072)	(0.026)	(0.007)	(0.041)		(0.073)	(0.032)	(0.007)	(0.038)		(0.073)	(0.041)	(0.006)	(0.035)	
$s_{n,t+1}$	0.431***	0.885***	-0.000	-0.129***	0.814	0.396***	0.834***	0.003	-0.107**	0.719	0.328***	0.754***	0.008	-0.031	0.557
	(0.078)	(0.028)	(0.009)	(0.047)		(0.078)	(0.035)	(0.008)	(0.045)		(0.080)	(0.046)	(0.007)	(0.042)	
r_{t+1}^i	0.139	-0.366**	0.515***	-0.106	0.324	0.124	-0.215	0.545***	0.042	0.313	0.137	-0.018	0.557***	0.132	0.400
	(0.284)	(0.150)	(0.077)	(0.217)		(0.282)	(0.143)	(0.070)	(0.220)		(0.283)	(0.169)	(0.068)	(0.219)	
rb_{t+1}	-0.382***	0.096***	-0.010	0.974***	0.711	-0.369***	0.157***	-0.012**	0.978***	0.726	-0.315***	0.272***	-0.018***	0.931***	0.756
	(0.070)	(0.026)	(0.007)	(0.039)		(0.071)	(0.032)	(0.006)	(0.037)		(0.071)	(0.041)	(0.006)	(0.033)	
	1985:1 – 2007:7														
$\Delta y_{1,t+1}$	-0.443***	0.087***	-0.015	0.098**	0.215	-0.433***	0.148***	-0.013	0.112**	0.249	-0.374***	0.279***	-0.015	0.075*	0.334
	(0.075)	(0.032)	(0.012)	(0.048)		(0.073)	(0.032)	(0.011)	(0.045)		(0.076)	(0.046)	(0.010)	(0.038)	
$s_{n,t+1}$	0.431***	0.886***	0.007	-0.121**	0.802	0.394***	0.830***	0.004	-0.110**	0.708	0.323***	0.736***	0.003	-0.038	0.523
	(0.081)	(0.036)	(0.015)	(0.055)		(0.082)	(0.043)	(0.014)	(0.054)		(0.084)	(0.053)	(0.012)	(0.048)	
r_{t+1}^i	0.242	-0.488***	0.399***	-0.358*	0.266	0.204	-0.416***	0.434***	-0.215	0.250	0.157	-0.281	0.465***	-0.032	0.236
	(0.265)	(0.140)	(0.083)	(0.195)		(0.265)	(0.155)	(0.081)	(0.193)		(0.266)	(0.183)	(0.080)	(0.194)	
rb_{t+1}	-0.401***	0.098***	-0.015	0.971***	0.687	-0.390***	0.156***	-0.015	0.982***	0.700	-0.332***	0.278***	-0.017*	0.940***	0.733
	(0.074)	(0.032)	(0.012)	(0.046)		(0.075)	(0.038)	(0.011)	(0.043)		(0.075)	(0.046)	(0.010)	(0.037)	

Notes: This table reports the estimated parameters of the benchmark VAR(1) model shown in Equation (2.3) for 10-, 5- and 2-year bonds. The state vector contains the first difference of the 1-month Treasury bill rate (Δy_1), the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill (s_n), the real interest rate (r^i) and the relative bill rate (*rb*). All variables are expressed in percentages per annum on continuously compounded basis. The upper panel of the table provides the full sample (1985:1 – 2014:2) estimates, while the pre-crisis period (1985:1 – 2007:7) estimates are shown in the lower panel. Heteroscedasticity and autocorrelation-consistent standard errors are shown in parentheses. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table 2.1 reports the estimated VAR(1) coefficients for the full and pre-crisis sample periods for three alternative models that only differ in terms of the zero-coupon bond yield used to calculate the long-short spread (10-, 5- or 2-year yields). Heteroscedasticity and autocorrelation-consistent standard errors are shown in the parentheses. The results can be summarised as follows.

Firstly, the one-month ahead forecasting power of the VAR is quite reasonable. The highest R^2 values are recorded in the spread equation, ranging from 56% to 81% in the full sample estimations and from 52% to 80% in the pre-crisis sample. The R^2 value decreases monotonically as the maturity of a longer-term bond decreases. With respect to the relative bill rate, the coefficient of multiple determination is also large and it is very similar across the three maturities. The full sample R^2 lies between 71% - 76%, while the pre-crisis values are within 69% - 73% range. In the remaining two equations, the R^2 value is lower but still reasonably high. For the first difference in the bill rate, it varies between 20% - 32% and 22% - 33% in the full and pre-crisis sample periods, respectively. It also increases in magnitude as the maturity of a longer-term bond, used to calculate the term spread, shortens. The R^2 reported for the real interest rate equation falls within the range of 32% - 40% in the full-sample. The highest value is associated with the 2-year bonds. The R^2 declines somewhat in the pre-crisis sample, ranging from 24% to 27% across the three models.

The change in the nominal short-term rate is predicted by its own lag, the lagged long-short spread and the lagged relative bill rate. The term spread parameter increases as one moves from the model with the 10-year bonds to 2-year bonds. Other coefficients in the first equation seem to be quite similar across the maturities. The long-short spread is highly persistent with its autoregressive coefficient being close to 0.90 across the VAR models, although it is slightly smaller in the case of the 2-year bonds. In addition, the spread can be forecast by the lagged relative bill rate, albeit not for the 2-year bonds, and the lagged change in the nominal short-term rate. The change in the bill rate has a somewhat smaller weight in the model with the shortest maturity bonds. The real interest rate typically follows an AR(1) process with the coefficient around 0.52 – 0.56 in the full sample estimation and between 0.40 – 0.47 prior to the crisis. The lagged spread generally helps to forecast the real rate in the case of the 10-year and 5-year bonds. The relative bill rate is forecast by its own lag with the coefficient estimate being around 0.94 – 0.98, the lagged spread and the lagged change in the nominal short-term rate. The coefficient on the term spread increases as maturity of a longer-term bond decreases, while the coefficient on the change in the short-term rate remains largely similar across the models.Regarding the magnitude, sign and statistical significance of the estimated coefficients, the findings in Table 2.1 are broadly in line with Campbell and Ammer (1993).

Overall, there are no substantial quantitative and qualitative changes in the VAR estimation results across the full and pre-crisis samples. This indicates that the dynamics of the system are not significantly affected by the financial crisis period. Finally, the estimated VARs are dynamically stable since no root lies outside the unit circle.[108]

2.7.2 Variance decomposition results

The results of the variance decomposition in Equation (2.2) for the 10-, 5- and 2-year bonds are shown in Table 2.2. The first six rows present the variances of the three components of unexpected excess bond returns and the covariance terms between these components normalised by the total variance of unexpected returns. In addition, the R^2 statistics are reported from the univariate regressions of unexpected excess returns on each of the estimated components in turn.

The key finding in Table 2.2 is that across different maturities news about future inflation is the dominant factor in explaining the variation of Treasury bond returns. With respect to the full sample, the variance decomposition attributes 83% of the variance of unexpected excess returns on the 10-year bonds to the variance of inflation news. The share of the total variance explained by the inflation news component increases as maturity decreases and amounts to 161% in the case of the 2-year bonds. It should be noted that both the volatility of inflation news and that of unexpected excess returns decrease as one moves from longer-term to shorter-term bonds, but the latter's decrease is more pronounced.[109] Hence, the ratio of the volatility of inflation news to the volatility of unexpected excess Treasury returns is higher in the case of shorter maturities. [110] The dominant role of inflation in explaining the bond market movements is also highlighted by the high R^2 values in the regressions of returns innovations on the inflation news term. Finally, the inflation component is also always statistically significant.

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[108] Note also that Augmented Dickey-Fuller and Phillips Perron unit root test results indicate that all state variables are stationary (see Table C2.1-Panel B in Appendix C).

[109] The standard deviation of unexpected excess Treasury bond returns declines from 35.08, in the case of the 10-year bonds, to 18.18 and 6.69 for the 5- and 2-year bonds, respectively. The corresponding figures for inflation news are 32.02, 19.14 and 8.49.

[110] This finding is also in line with several studies that carry out the variance decomposition for the one-month interest rate and show that inflation news is much more volatile than the short rate itself (Campbell and Ammer, 1993; Bar and Pesaran, 1997). As compared to the results for long-term bond returns decomposition, this implies that the revisions in expectations about inflation tend to revert but not fast enough to leave the long end of the yield curve unaffected.

Table 2.2: Variance decomposition for excess bond returns

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var\left(\tilde{x}_{\pi}\right)$	0.833*** (0.264)	0.799*** (0.243)	1.108** (0.430)	1.116** (0.447)	1.607** (0.763)	1.675** (0.700)
$2Cov\left(\tilde{x}_{\pi},\tilde{x}_{r^i}\right)$	-0.087 (0.083)	-0.085 (0.087)	-0.189 (0.189)	-0.225 (0.191)	-0.674 (0.698)	-0.673 (0.595)
$2Cov\left(\tilde{x}_{\pi},\tilde{x}_{x}\right)$	0.046 (0.285)	0.148 (0.200)	-0.098 (0.352)	0.015 (0.331)	-0.553 (0.448)	-0.314 (0.421)
$Var\left(\tilde{x}_{r^i}\right)$	0.018* (0.010)	0.017 (0.012)	0.038 (0.023)	0.035 (0.024)	0.266* (0.147)	0.143 (0.111)
$2Cov\left(\tilde{x}_{r^i},\tilde{x}_{x}\right)$	-0.068 (0.056)	-0.102* (0.058)	0.025 (0.067)	-0.052 (0.069)	0.252 (0.171)	0.081 (0.128)
$Var\left(\tilde{x}_{x}\right)$	0.258 (0.168)	0.223 (0.177)	0.116 (0.101)	0.112 (0.109)	0.102 (0.074)	0.088 (0.070)
$R^2\left(\tilde{x}_{\pi}\right)$	0.793*** (0.154)	0.863*** (0.140)	0.839*** (0.113)	0.915*** (0.083)	0.614*** (0.163)	0.834*** (0.102)
$R^2\left(\tilde{x}_{r^i}\right)$	0.199 (0.145)	0.334** (0.152)	0.050 (0.133)	0.314 (0.202)	0.011 (0.087)	0.164 (0.252)
$R^2\left(\tilde{x}_{x}\right)$	0.236 (0.202)	0.271 (0.203)	0.054 (0.221)	0.078 (0.268)	0.023 (0.149)	0.009 (0.103)

Notes: This table reports the variance decomposition of unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_{x}$) and the covariances between these three components. News components are extracted from a VAR(1) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample (1985:1 – 2014:2) and pre-crisis period (1985:1 – 2007:7) results, respectively. R^2 values are obtained from the regressions of unexpected excess returns on each news component in turn. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

On the other hand, the estimates of the relative variance terms of news about risk (term) premium are substantially smaller, especially for the 5- and 2-year bonds, and are never statistically significant. Nevertheless, it still explains about 26% of the total excess return volatility for the 10-year maturity. This is consistent with longer-term bonds being viewed by market participants as more risky securities. Meanwhile, the relative variance of real interest rate news is less than 5% for longer maturities; however, it becomes greater for the 2-year bonds and is approximately 27% for the full sample estimation.[111] Nevertheless, this term is statistically insignificant at conventional significance levels. The finding that real rate news is not relevant for longer-term bonds but its importance increases somewhat for shorter-term bonds indicates that the revisions in expectations about real rate decay over the long period even though it may be more variable in the short run. In the long run, investors seem to be accepting the Fisher hypothesis (Campbell and Ammer, 1993; Barr and Pesaran, 1997).

With respect to the covariance terms, they typically play only a minor and statistically insignificant role in the decomposition of returns variability. Nevertheless, for the 2-year bonds, large and negative covariances between news about inflation and future excess returns and between inflation and real rate news partially offset the large relative share of the inflation component in this case. Thus, when investors revise their expectations about future inflation upwards, they also appear to be expecting lower future excess returns and lower real interest rates.

When the financial crisis and its aftermath are excluded from the VAR estimation sample, the obtained variance decompositions remain broadly unchanged. This finding is consistent with the fact that the VAR estimation results in Table 2.1 do not indicate significant changes across the two samples in the predictability of the components of excess bond returns.[112]

Overall, news about future inflation appear to be very important for the bond market in the US as it explains a large share of the variability in unexpected excess bond returns across the term structure. Also, this result indicates that bond yields may have some ability to forecast inflation over the long and shorter horizons. The importance of inflation news is consistent with the previous evidence for the US over sample periods that include the highly inflationary 1970s and the early 1980s (Campbell and Ammer, 1993; Engsted and Tanggaard, 2007; Bredin, Hyde and O'Reilly, 2010). Thus, revisions in inflation

[111] However, the R^2 values do not confirm such tendency. This may be due to the fact that the coefficients of determination are not orthogonalised.

[112] The R^2 statistics from the VAR model equations for the change in the nominal short-term risk-free rate and the term spread remain fairly stable when the financial crisis and its aftermath are removed from the sample.

expectations maintained their dominant influence over the Treasury bond market even in the era of lower inflation since the mid-1980s.

2.7.3 Monetary policy effects on unexpected excess returns and their components

Tables 2.3-2.6 report the estimates of the impact of monetary policy actions on unexpected excess Treasury bond returns and their components over the full and pre-crisis sample periods.[113] The results in Tables 2.3 and 2.4 are based on the interest rate measures of monetary policy (actual and unexpected change in the FFR, respectively).

Table 2.3: Impact of monetary policy on excess bond returns – FFR change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-20.876***	-20.440***	-15.588***	-13.840***	-8.507***	-7.871***
	(4.662)	(4.411)	(2.969)	(2.801)	(1.349)	(1.305)
$\tilde{x}_{r^i}^{MP}$	0.940	1.553	-1.514	-0.714	-2.094	-1.293
	(1.829)	(1.663)	(1.257)	(1.078)	(1.593)	(1.374)
$\tilde{x}_\pi^{MP}$	36.292***	32.902***	23.033***	19.671***	13.111***	11.192***
	(8.331)	(7.532)	(3.668)	(3.332)	(1.947)	(1.661)
$\tilde{x}_x^{MP}$	-16.356**	-14.014**	-5.932*	-5.117*	-2.509**	-2.028*
	(7.833)	(6.878)	(3.544)	(3.089)	(1.265)	(1.126)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). The news components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample (1985:1 – 2014:2) and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

[113] Note that the VAR model that generates excess bond returns innovations and the associated news components is estimated over the full sample period. The use of a longer sample should improve the precision of the estimates. Nevertheless, if the pre-crisis monetary policy regressions are estimated using returns and news components extracted from the VAR estimated with pre-crisis data only, the results (not reported to preserve space) are very similar to those reported in Tables 2.3-2.6.

Table 2.4: Impact of monetary policy on excess bond returns – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-24.318*** (1.963)	-54.002*** (3.511)	-25.568*** (1.341)	-34.730*** (1.799)	-17.340*** (0.712)	-19.219*** (0.732)
$\tilde{x}_{r^i}^{MP}$	-1.780** (0.863)	-2.367 (1.444)	-0.564 (1.616)	-0.729 (1.943)	2.041 (3.305)	1.898 (3.461)
$\tilde{x}_\pi^{MP}$	16.704*** (3.687)	44.209*** (6.009)	24.037*** (5.145)	33.432*** (5.804)	16.314*** (4.713)	18.450*** (4.726)
$\tilde{x}_x^{MP}$	9.394** (3.951)	12.159* (6.963)	2.096 (4.521)	2.028 (5.736)	-1.016 (2.551)	-1.129 (2.727)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on the FFR futures, the full sample that is used for the estimations of monetary policy effects commences in 1989:2. See also Table 2.3 notes.

The first main finding is that monetary policy actions significantly affect the bond market across all three maturities and across both sample periods. In the case of actual changes in the funds rate, the impact is smaller. It is expected since this proxy of policy actions also includes the expected component of a change in the rate that should not affect financial markets. Monetary easing (FFR cuts) is associated with higher contemporaneous unexpected excess returns. For instance, the response of unexpected excess returns on the 10-year bonds to a 1-percentage-point (1%) surprise FFR cut is approximately 24% (around 2% on a monthly basis) in the full sample. The change in bond returns is economically relevant, although it tends to decrease with the maturity of bonds.[114] In the case of the 2-year bonds, the full sample response coefficient is 17.34% (1.45% monthly). The estimated reaction coefficinets are similar in magnitude to the response of unexpected stock and bond returns as reported in similar studies (Valckx, 2004; Maio, 2014). Higher bond returns imply lower longer-term yields indicating that while central banks control a short-term policy rate directly their actions also have implications for the whole term structure of interest rates. This provides support for the interest rate channel of monetary policy transmission mechanism.

The second main result is that the effect of monetary policy actions on the bond market is largely explained through the inflation news channel regardless of the interest

[114] As discussed in the literature review, it is typically found that the monetary policy impact declines for longer-term interest rates. The approach here, instead, considers unexpected excess bond returns. Moreover, there are various possible explanations for the significant reaction at the long end of the bond market. For instance, Rolley and Sellon (1995) point out that if policy actions are seen as relatively permanent or as the first in a series of future actions, the response of long-term rates may be larger than the response of short-term rates. Finally, longer-term bonds are more exposed to uncertainty about future developments, thus, returns are more volatile.

rate proxy used. Specifically, it is found that the key driver of the positive bond returns in response to expansionary policy is its negative effect on inflation expectations.[115] In the full sample, the monetary policy impact on revisions in expectations about future inflation explains 16.70 percentage points of the response of returns on the 10-year bonds to a monetary policy surprise. This amounts to more than a half of the total policy effect on long-term bonds. With respect to the 5- and 2-year maturities, a change in inflation expectations following a policy surprise accounts for almost entire bond returns reaction. In the case of raw funds rate changes, inflation news component is even more responsive than returns themselves. As with the bond returns, the response of their main component, inflation news, tends to increase in magnitude, albeit not monotonically, as the maturity increases.

While expansionary monetary policy exerts a large and statistically significant effect on inflation expectations, the impact on news about expected excess bond returns (term premium) is typically smaller, especially for the shorter-term bonds. In the case of the 10-year bonds, the response of future expected excess bond returns to a surprise decrease in the funds rate is 9.39%, i.e. less than a half of the total returns response. It is also statistically significant at the 5% level. Thus, unexpected policy expansion is associated with a decrease in the expected term premium. This is consistent with the findings for the stock market in Bernanke and Kuttner (2005). Nevertheless, the sign of the parameter on risk premium news differs across the two interest rate measures. Using actual FFR changes, the positive effect of monetary easing on expected excess returns is outweighed by the negative effect on inflation expectations, so that the total effect on bond returns is positive.[116] Moreover, both actual and unexpected rate changes have a much smaller and typically insignificant impact on news about future excess returns for the 5- and 2-year bonds. Finally, the response of revisions in real interest rate expectations to both actual and surprise changes in the policy rate is very small and typically statistically insignificant across the three maturities.

[115] Bredin, Hyde and O'Reilly (2010) report a similar finding for the UK. It could possibly be explained through the ability of a central bank to control inflation. If surprise policy expansion leads to lower inflation expectations, it may be indicative of the lack of credibility with respect to fighting deflationary pressures. Also, when monetary easing takes place during the periods of financial turmoil, it may reinforce flight to safety and therefore increase the price of Treasuries (Kontonikas, MacDonald and Saggu; Goyenko and Ukhov, 2009). Alternatively, such a bond market's response could be explained by the monetary policy impact on financial market sentiment. If expansionary policy actions induce the pessimism among investors regarding the future state of economy, bond yields are likely to fall.

[116] Re-arranging the dynamic identity shown in Equation (2.12), it can be seen that the total monetary policy effect on unexpected excess bond returns must be equal to the negative sum of the effects on inflation, real interest rate and risk premium news.

With respect to the pre-crisis sample period, the results are very similar to those for the full sample analysis if actual FFR rate changes are used to measure policy actions. Nevertheless, there are some notable changes in magnitudes of response coefficients across the two sample periods if surprise policy rate changes are employed. This may be explained by several positive monetary policy surprises in the late 2008 that coincide with increasing asset prices and decreasing bond yields. This could have reduced the overall response of the returns and its components in the full sample as compared to the pre-crisis period. On the other hand, the findings are qualitatively the same.

Table 2.5: Impact of monetary policy on excess bond returns – MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.779*** (0.271)	-0.913 (1.076)	0.796*** (0.095)	0.749 (0.531)	0.337*** (0.034)	0.491** (0.207)
$\tilde{x}_{r^i}^{MP}$	0.167* (0.095)	-0.552** (0.251)	0.254*** (0.085)	-0.463** (0.189)	0.273*** (0.090)	-0.456** (0.184)
$\tilde{x}_\pi^{MP}$	-2.136*** (0.354)	0.725 (1.495)	-1.485*** (0.198)	-0.255 (0.748)	-0.787*** (0.117)	0.059 (0.361)
$\tilde{x}_x^{MP}$	1.191** (0.541)	0.740 (0.786)	0.434* (0.224)	-0.031 (0.331)	0.177*** (0.066)	-0.094 (0.117)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table 2.3 notes.

Table 2.6: Impact of monetary policy on excess bond returns – Unexpected MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	1.054*** (0.102)	1.232** (0.537)	1.187*** (0.048)	1.066*** (0.328)	0.510*** (0.018)	0.332** (0.147)
$\tilde{x}_{r^i}^{MP}$	0.301** (0.123)	0.017 (0.251)	0.340** (0.138)	-0.293 (0.209)	0.305* (0.159)	-0.520*** (0.184)
$\tilde{x}_\pi^{MP}$	-2.160*** (0.596)	1.650 (1.233)	-1.862*** (0.407)	0.383 (0.719)	-1.007*** (0.234)	0.532 (0.322)
$\tilde{x}_x^{MP}$	0.805 (0.533)	-2.899*** (1.102)	0.335 (0.294)	-1.156** (0.511)	0.193* (0.103)	-0.343** (0.151)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table 2.3 notes.

The findings in Tables 2.5 and 2.6 are based on two policy indicators that are related to the (log) monetary base (actual and unexpected change, respectively). The focus here is on the full sample estimation results since the pre-crisis sample excludes the period when quantity-based indicators became strongly active due to the unconventional policies adopted by the Fed. The main insights that can be identified using the interest-rate-based measures remain overall valid in the full sample estimations with the quantity-based measures. For instance, an unexpected 1% (percentage point) increase in the monthly growth of monetary base leads to 1.1% (0.09% on a monthly basis) unexpected excess returns on the 10-year Treasuries. In November 2008, the unexpected growth in monetary base as measured in Equation (2.14) was 10.67%. This figure is associated with 0.96% response in unexpected monthly excess bond returns for the 10-year bonds. The impact is the smallest for the 2-year bonds, in line with the results in Tables 2.3 – 2.4. The full sample estimated response parameters are somewhat smaller in magnitude when actual change in monetary base is used. This again indicates that financial asset prices are more reactive to unexpected changes in monetary policy stance.

With respect to the news components, the positive effect of monetary easing (higher monetary base growth) on unexpected excess Treasury bond returns comes through downward revisions in inflation expectations, with the impact being even stronger than the total response of unexpected returns and also being generally stronger at longer maturities. For instance, the response of the inflation news component for the 10-year bonds to a surprise monetary base growth is -2.16% as compared to a 1.1% response of unexpected returns. For all maturities, the effect on returns due the negative impact on inflation news is dampened by the positive policy effects on other two news components. The full sample results indicate that money growth significantly affects real interest rate expectations, whereas the impact on risk premium news tends to be statistically significant only when actual changes in the (log) monetary base are used. The response of the expected excess returns component is greater in the case of longer maturities, meanwhile the real interest rate component tends to be more responsive for shorter maturities. Hence, monetary policy expansion through a higher monetary base growth rate is associated with upward revisions in expectations about future real interest rate and excess bond returns, i.e. risk premium. This finding likely reflects the impact of the financial crisis since the pre-crisis sample estimates are both negative instead. Nevertheless, the positive effect of monetary easing on expected future excess returns and real interest rates is more than compensated by the negative impact on inflation expectations leading to an overall increase in unexpected bond returns.

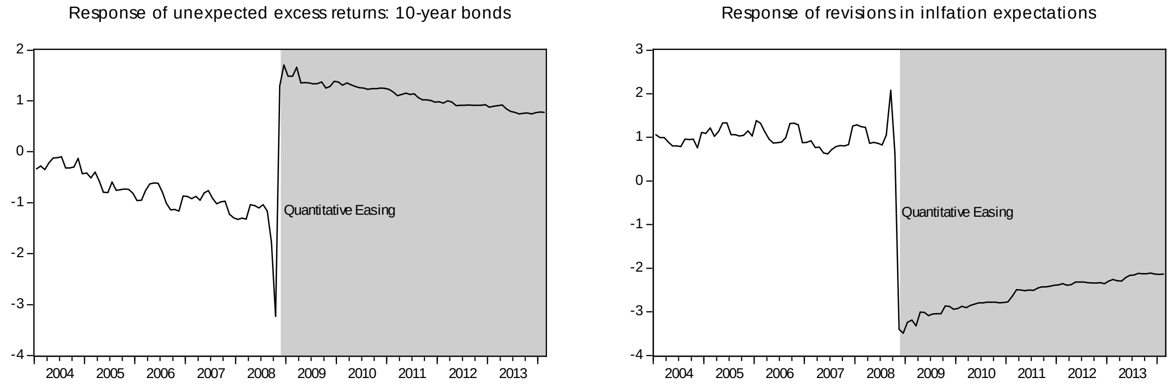
Comparing the full sample with the pre-crisis results from the quantity-based measures of monetary policy, it becomes apparent that the former largely reflect developments that occurred during the financial crisis. Following the collapse of Lehman

Brothers in September 2008, inflation expectations sharply deteriorated in line with worsening economic outlook (Campbell, Shiller and Viceira, 2009). By the autumn of 2008, inflation became strongly negative recording a sample minimum of -1.8% (month-on-month) in November 2008. The nominal short-term interest rate fell to almost zero, thereby pushing up the ex post real interest rate to highly positive values. At the same time, the Fed significantly expanded the pace of monetary easing, both in the conventional and unconventional sense. The federal funds rate declined by 160 basis points from September to November and the monetary base growth rate recorded historical highs due to the heavy usage of non-sterilised Fed liquidity facilities. Figures 2.3 and 2.4, which plot the recursive estimates of the impact of actual and unexpected (log) monetary base changes on unexpected excess Treasury bond returns and inflation news, also suggest that an important structural shift took place in autumn 2008. Following the unprecedented expansion in the monetary base and the announcement of QE1, the relationship between money growth and bond returns tends to increase in magnitude, while the impact on inflation expectations becomes strongly negative. The response parameters exhibit a tendency to become smaller in size after the initial shock, suggesting that further rounds of QE may not have been as influential as the first one.

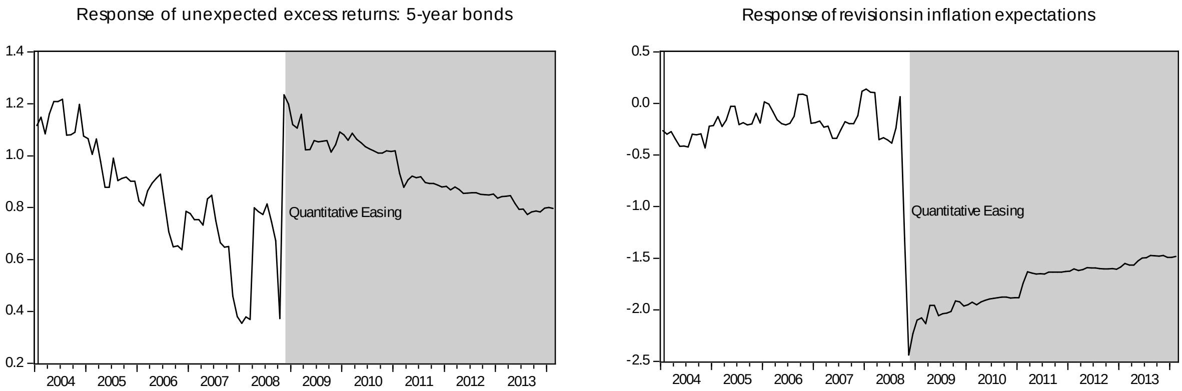
Summarising the main results, the positive effect of monetary easing on the Treasury bond market is principally due to falls in inflation expectations. Moreover, the results reported in this chapter are overall not supportive of the portfolio balance mechanism, according to which, monetary easing via an expansion of the Fed’s balance sheet should increase current period bond returns primarily through downward adjustments in expected excess returns, i.e. the risk (term) premium.

Figure 2.3: Recursive estimates of MB change impact

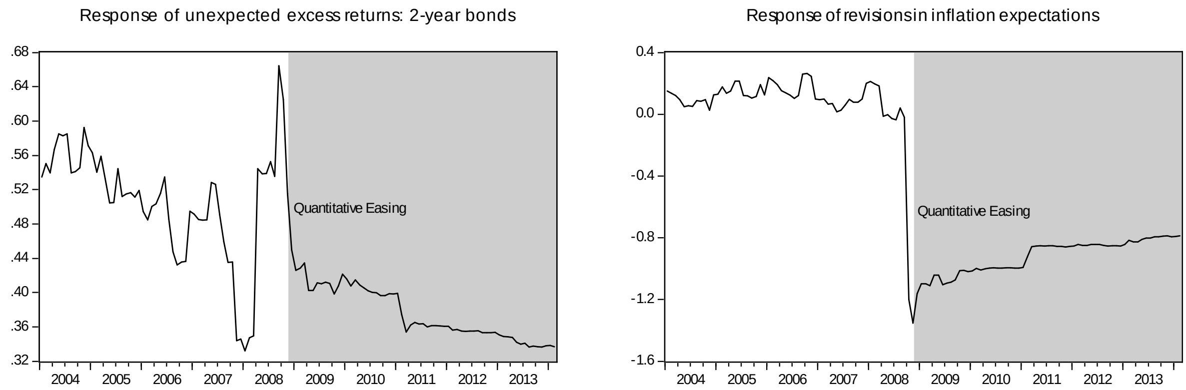
Panel A: 10-year bonds



Panel B: 5-year bonds



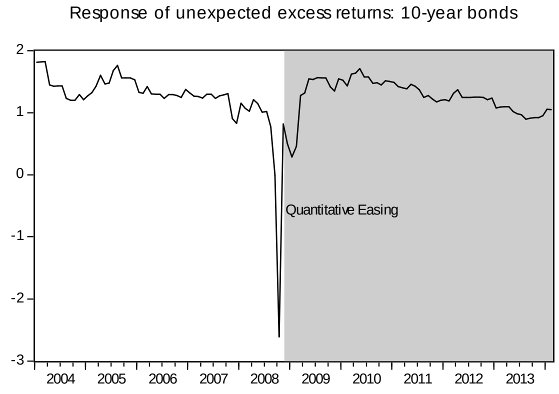
Panel C: 2-year bonds

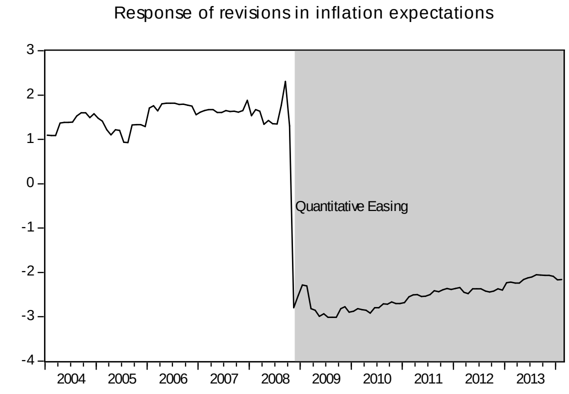


Notes: This figure plots the recursive estimates of the response parameters of unexpected excess Treasury bond returns and the corresponding inflation news component to actual changes in log monetary base (MB). Panel A refers to 10-year bonds, Panel B to 5-year bonds and Panel C to 2-year bonds. The initial sample of the recursive estimation is 1985:1 – 1995:1 and then one month is added at each step. The shaded area denotes the period of quantitative easing, starting from the announcement of QE1 (2008:11).

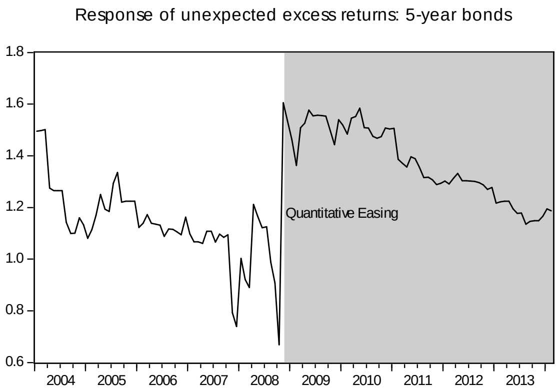
Figure 2.4: Recursive estimates of unexpected MB change impact

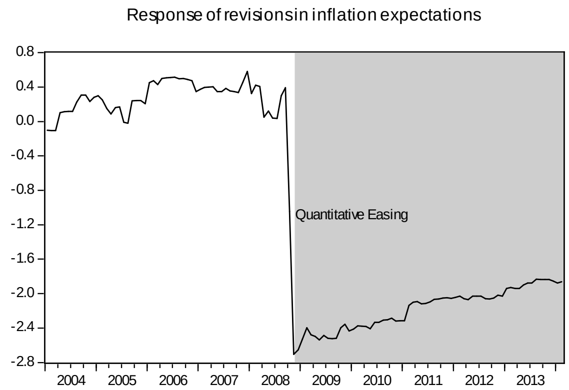
Panel A: 10-year bonds



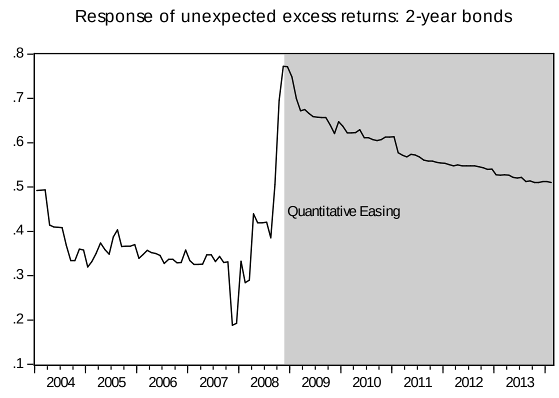


Panel B: 5-year bonds





Panel C: 2-year bonds



Notes: This figure plots the recursive estimates of the response parameters of unexpected excess Treasury bond returns and the corresponding inflation news component to unexpected changes in log monetary base. See also Figure 2.3 notes.

2.8 Robustness analysis

The robustness of the empirical findings is examined in a number of ways. It is shown that overall the baseline results reported in Section 2.7 are not sensitive to these changes. First, monetary policy effects are estimated over an alternative sample period. Second, alternative state vector specifications for the underlying VAR model are used. Third, an alternative interest-rate-based policy indicator that accounts for the Fed's private information is employed. Fourth, higher-order VAR models are considered. Fifth, the model that is used to extract monetary base growth surprises is modified. Finally, alternative quantity-based monetary policy indicators are considered. The results are contained in Appendix C.

2.8.1 Alternative sample period

In the early 1990s, the Fed's decisions to cut the policy target rate may have reflected an endogenous reaction to labour market conditions. Between June 1989 and September 1992 (the date of the last FFR cut associated with employment news), nearly half of the FOMC meetings coincided with the release of a worse-than-expected employment report (Bernanke and Kuttner, 2005). In this section, the sensitivity of the baseline findings is examined with respect to the exclusion of the pre-October 1992 period.[117] The results are presented in Tables C2.4 and C2.5 of the Appendix C. With respect to the 2-year bonds, they are qualitatively similar to the main findings, with the positive effect of monetary easing on bond returns being primarily explained by downward revisions in inflation expectations. Nevertheless, the magnitude of the related coefficients is reduced. Meanwhile, the results for the 5- and 10-year bonds are sensitive to the exclusion of the pre-October 1992 period. Specifically, the evidence of the significant bond market reaction to monetary policy shifts, explained through the inflation news channel, becomes overall weaker.[118]

In addition, an alternative sample period commencing in February 1994, when the Fed started to announce FFR target changes and reduced substantially the number of intermeeting policy rate changes, is also considered. The results (see Tables C2.6-C2.7) for

[117] Only conventional monetary policy measures are considered here since the results with respect to unconventional policies are driven by the end of the sample. For monetary base measures, the results (not reported to preserve space) are qualitatively the same.

[118] The puzzling full sample finding of a positive and statistically significant response of 10-year bonds returns to monetary tightening surprises is driven by the crisis period developments.

the 5- and 10-year bonds deteriorate further, while in the case of the 2-year bonds they remain broadly similar. The weaker bond market reaction to FFR shifts over the more recent period may be related to changes in the way that the Fed implements and communicates monetary policy (Fawley and Neely, 2014). These changes have enhanced transparency and enabled financial markets to form more accurate expectations regarding the policy rate, leading to overall smaller and less volatile target rate surprises over time.

2.8.2 Alternative state vector specifications

The benchmark VAR state vector includes the change in the nominal short-term risk-free rate, the term spread, the real interest rate, and the relative bill rate. In addition to interest rate variables, some studies find that macroeconomic factors and financial conditions indicators are helpful in predicting bond returns (Ang and Piazzesi, 2003; Ludvigson and Ng, 2009; Fricke and Menkhoff, 2014). Motivated by this evidence, it is examined whether the baseline findings are robust to incorporating measures of macro-financial conditions in the VAR state vector. The following variables are considered: the industrial production growth rate, unemployment rate, the Chicago Fed National Activity Index (CFNAI), and the Chicago Fed Adjusted National Financial Conditions Index (ANFCI).[119] CFNAI is a measure of overall economic activity, calculated as the weighted average of 85 monthly indicators of national economic activity. ANFCI isolates the component of financial conditions (in money markets, debt and equity markets, and the traditional and "shadow" banking systems) that is uncorrelated with economic conditions.

The variance decomposition results using the alternative state vectors are shown in Tables C2.8-C2.11 in Appendix C, while the corresponding monetary policy effects regressions are presented in Tables C2.12-C2.27. Overall, as in the benchmark case, inflation news is the major component of unexpected excess Treasury bond returns. Furthermore, in line with the baseline results, the positive effect of monetary easing on bond returns comes from a corresponding negative effect on inflation expectations. Thus, accounting for additional forecasting variables does not alter the conclusions from the main analysis.

2.8.3 Alternative interest rate-based policy measure

[119] Both CFNAI and ANFCI may provide useful information about current and future developments in economic and financial conditions. More details about the indices can be found at: https://www.chicagofed.org/publications/cfnai/index; https://www.chicagofed.org/publications/nfci/index.

If policy actions reveal some private information held by a central bank about the future state of the economy, the estimates of monetary policy effects on economic and financial variables may be biased. Romer and Romer (2004) propose an alternative way to identify monetary policy shocks that takes into account the central bank's response to expected economic conditions.[120] The results presented in Table C2.28 are obtained using Romer and Romer's monetary policy shocks. The conclusions that can be drawn are similar to those from the baseline findings in Tables 2.3 and 2.4, since bond returns respond positively to monetary easing and inflation expectations play a key role in explaining this reaction.

2.8.4 Higher order VARs

The benchmark VAR model is a first-order VAR. In order to examine whether a more complex dynamic structure affects the baseline results, higher order VARs are estimated (Barr and Pesaran, 1997; Maio, 2014). The variance decomposition and monetary policy effects results in Tables C2.29 and C2.30-C2.33, respectively, are based upon a third-order VAR model. They indicate that the main conclusions about the role of inflation news in the variance decomposition, as well as the relationship between bond returns and monetary policy, and are not affected by parsimony in the VAR order. Similar insights are provided by the VAR(6) model with the results summarised in Tables C2.34 and C2.35-C2.38.

2.8.5 Alternative models for monetary base growth surprises

The monetary base growth surprises in the baseline analysis are obtained as residuals from the regression of monetary base growth on its own lags and the lags of unemployment measure. Following the previous work by Cover (1992), monetary base growth is modelled using three alternative specifications with respect to the set of explanatory variables. Firstly, the lags of monetary base growth and unemployment measure are complemented with the lags of the Chicago Fed Adjusted National Financial Conditions Index in order to take into account a posibble policy response to financial

[120] The calculation of Romer and Romer's (2004) monetary policy shocks involves two steps. First, intended federal funds rate changes around the FOMC meetings are identified. Second, the intended funds rate changes are regressed on the internal FOMC forecasts for inflation and real economic activity, i.e. the Greenbook forecasts, around the dates of these forecasts; see Equation (1) in Romer and Romer (2004). Residuals from that regression represent monetary policy shocks. To obtain these shocks, the STATA code and data provided by Wieland and Yang (2015) is used.

conditions. Also, the lags of monetary base growth are then combined with either the lags of industrial production growth, or the lags of industrial production growth and the first difference of the 3-month Treasury bill rate. The estimates of monetary policy effects using these alternative measures of monetary base growth surprises are presented in Tables C2.39 - C2.41 in Appendix C. They are overall similar to the benchmark results. The positive bond market response to monetary easing is mainly explained by downward revisions in inflation expectations.

2.8.6 Alternative quantity-based monetary policy indicator

The large increase in the total reserves, since the end of 2008, made them the dominant component of the monetary base. Motivated by this development, two additional quantity-based measures of monetary policy are considered: actual and unexpected changes in (log) the total reserves. As with monetary base growth surprises, the latter are obtained as residuals from the regression of the total reserves growth on its own lags and the lags of unemployment. The results from the monetary policy regressions with the total reserves as a quantity-based indicator are shown in Tables C2.42 and C2.43 in the Appendix C. The main conclusions from the baseline analysis remain valid since monetary policy shifts have a significant effect on bond market performance and inflation news is typically the main component of bond returns that is affected.

2.9 Conclusion

Following the recent financial crisis and the actions taken by the Fed, the analysis of the sources of variation in the bond market and the role of monetary policy came to the focus of academics, investors and policymakers. This chapter extends the analysis of Campbell and Ammer (1993) to investigate the sources of variation in Treasury bond returns across different maturities. This framework combines a dynamic accounting identity with a VAR time-series econometric model to decompose unexpected excess bond returns into the revisions in expectations (news) about future excess returns, inflation and real interest rates. Furthermore, the extension of Campbell and Ammer's framework by Bernanke and Kuttner's (2005) is modified to obtain insights to the sources of the bond market response to monetary policy actions. Using this approach, the impact of actual and unexpected changes in monetary policy indicators on bond returns and their components is estimated. The federal funds rate-based indicators are used to capture conventional monetary policy, whereas shifts in the monetary base are employed to capture the unconventional dimensions of monetary policy during the crisis and its aftermath.

The variance decomposition results show that news about future inflation constitutes the largest component of unexpected excess Treasury bond returns, while the contribution of news about future excess returns (risk premium) and real interest rate news is typically negligible. Hence, the findings confirm and update previous empirical evidence about the importance of inflation news for longer-term bonds by showing that it maintained the dominant influence during the era of lower inflation that commenced in the mid-1980s. Moreover, this chapter completes the picture by providing new evidence which shows that inflation news also dominates the variance decomposition of medium- and shorter-term bonds.

With respect to the impact of monetary policy actions, the results generally indicate that monetary easing is associated with higher bond returns. Nevertheless, the effect of interest rate-based policy measures on bond returns has become weaker over the more recent period possibly reflecting changes, ever since the mid-1990s, in the way that the Fed implements and communicates monetary policy. In the case of the quantity-based monetary policy indicators, the bond market response largely reflects developments that occurred at the peak of the financial crisis in autumn 2008. As to why the bond market responds in this manner, the results highlight the role of inflation news. The results show that the positive effect of monetary easing on bond returns mainly comes from the corresponding negative effect on inflation expectations. On the other hand, the evidence in favour of the portfolio balance mechanism’s prediction of a strong role for the risk premium within the context of expanding Fed’s balance sheet is rather elusive.

References for Chapter 2

Akhtar, M.A., 1995. Monetary policy and long-term interest rates: A survey of empirical literature. *Contemporary Economic Policy*, 13(3), p.110-130.

Ammer, J., and Mei, J., 1996. Measuring international economic linkages with stock market data. *The Journal of Finance*, 51(5), p.1743-1763.

Ang, A., and Piazzesi, M., 2003. A no-arbitrage vector autoregression of term structure dynamics with macroeconomic and latent variables. *Journal of Monetary Economics*, 50(4), p.745-787.

Ang, A., Boivin, J., Dong, S., and Loo-Kung, R., 2011. Monetary policy shifts and the term structure. *Review of Economic Studies*, 78(2), p.429-457.

Balke, N.S., and Emery, K.M., 1994. Understanding the price puzzle. *Federal Reserve Bank of Dallas Economic Review*, 4, p.15-26.

Bansal, R., and Shaliastovich, I., 2013. A long-run risks explanation of predictability puzzles in bond and currency markets. *The Review of Financial Studies*, 26(1), p. 1-33.

Barakchian, S.M., and Crowe, C., 2013. Monetary policy matters: Evidence from new shocks data. *Journal of Monetary Economics*, 60(8), p.950-966.

Barr, D., and Pesaran, B., 1997. An assessment of the relative importance of real interest rates, inflation, and term premiums in determining the prices of real and nominal U.K. bonds. *The Review of Economics and Statistics*, 79(3), p.362-366.

Bauer, M.D., and Neely, C.J., 2014. International channels of the Fed's unconventional monetary policy. *Journal of International Money and Finance*, 44, p.24-46.

Bauer, M.D., and Rudebusch, G.D., 2014. The signalling channel for Federal Reserve bond purchases. *International Journal of Central Banking*, 10(3), p.233-289.

Baumeister, C., and Benati, L., 2013. Unconventional monetary policy and the great recession: Estimating the macroeconomic effects of a spread compression at the zero lower bound. *International Journal of Central Banking*, 9(2), p.165-212.

Beckworth, D., Moon, K.P., and Toles, J.H., 2012. Can monetary policy influence long-term interest rates? It depends. *Economic Inquiry*, 50(4), p.1080-1096.

Belke, A., Gros, D., and Osowski, T., 2016. Did quantitative easing affect interest rates outside the US? New evidence based on interest rate differentials. *CEPS Working Documents*, Working paper 416.

Bernanke and Mihov, 1998. Measuring monetary policy. *The Quarterly Journal of Economics*, 113(3), p.869-902.

Bernanke, B.S., and Blinder, A.S., 1992. The federal funds rate and the channels of monetary transmission. *The American Economic Review*, 82(4), p.901-921.

Bernanke, B.S., and Kuttner, B.N., 2005. What explains the stock market's reaction to Federal Reserve policy? *Journal of Finance*, 60(3), p.1221-1257.

Berument, H., and Froyen, R., 2009. Monetary policy and U.S. long-term interest rates: How close are the linkages? *Journal of Economics and Business*, 61, p.34-50.

Botshekan, M., Kraeussl, R., and Lucas, A., 2012. Cash flow and discount rate risk in up and down markets: what is actually priced? *Journal of Financial and Quantitative Analysis*, 47(6), p.1279-1301.

Bredin, D., Hyde, S., and O'Reilly, G., 2010. Monetary policy surprises and international bond markets. *Journal of International Money and Finance*, 29(6), p.988-1002.

Bredin, D., Hyde, S., O'Reilly, and G., Nitzsche, D., 2007. U.K. stock returns and the impact of domestic monetary policy shocks. *Journal of Business, Finance and Accounting*, 34(5-6), p.872–888.

Bredin, D., O'Reilly, G., and Stevenson, 2011. Monetary policy transmission and real estate investment trusts. *International Journal of Finance and Economics*, 16(1), p.92-102.

Breedon, F., 2012. A variance decomposition of index-linked bond returns. *Economics Letters*, 116(1), p.49-51.

Brunnermeier, M.K., 2009. Deciphering the liquidity and credit crunch 2007-2008. *Journal of Economic Perspectives*, 23(1), p.77-100.

Cahill, M.E., D'Amico, S., Li, C., and Sears, J.S., 2013. Duration risk versus local supply channel in Treasury yields: Evidence from the Federal Reserve's asset purchase announcements. *Federal Reserve Board Finance and Economics Discussion Series*, Working paper 2013-35.

Campbell, J. Y., and Ammer, J., 1993. What moves the stock and bond markets? A variance decomposition for long term asset returns. *The Journal of Finance,* 48(1), p.3-37.

Campbell, J.R., Evans, C.L, Fisher, J.D.M., and Justiniano, A., 2012. Macroeconomic effects of Federal Reserve forward guidance. *Brookings Papers on Economic Activity*, Spring Issue, p.1-80.

Campbell, J.Y., 1991. A variance decomposition for stock returns. *Economic Journal*, 101(405), p.157-179.

Campbell, J.Y., and Shiller, R.J., 1988. The dividend–price ratio and expectations of future dividends and discount factors. *The Review of Financial Studies*, 1(3), p.195-228.

Campbell, J.Y., and Shiller, R.J., 1991.Yield spreads and interest rate movements: A bird's eye view. *Review of Economic Studies*, 58(3), p.495-514.

Campbell, J.Y., Lo, A.W., and MacKinlay, A.C., 1997. *The econometrics of financial markets*. New Jersey: Princeton University Press.

Campbell, J.Y., Pflueger, C., and Viceira, L.M., 2015. Monetary policy drivers of bond and equity risks. Working paper (unpublished).

Campbell, J.Y., Shiller, R.J., and Viceira, L.M., 2009. Understanding inflation-indexed bond markets. *NBER Working Paper Series*, Working paper 15014.

Cenedese, G., and Mallucci, E., 2016. What moves international stock and bond markets?. *Journal of International Money and Finance*, 60, p.94-113.

Cenesizoglu, T., and Essid, B., 2012. The effect of monetary policy on credit spreads. *The Journal of Financial Research*, 35(4), p.581-613.

Chen, L., and Zhao, X., 2009. Return decomposition. *The Review of Financial Studies*, 22(12), p.5213-5246.

Chen, Q., Filardo, A., He, D., and Zhu, F., 2015. Financial crisis, US unconventional monetary policy and international spillovers. *Journal of International Money and Finance*, forthcoming.

Chen, S., 2007. Does monetary policy have asymmetric effects on stock returns? *Journal of Money, Credit and Banking*, 39(2-3), p.667-688.

Chernov, M, and Mueller, P., 2012. The term structure of inflation expectations. *Journal of Financial Economics*, 106(2), p.367-394.

Christensen, J.H.E., and Rudebusch, G.D., 2012. The response of interest rates to US and UK quantitative easing. *Economic Journal*, 122(564), p.385-414.

Christopoulos, D.K., and Leon-Ledesma, M.A., 2007. A long-run non-linear approach to the Fisher effect. *Journal of Money, Credit, and Banking*, 39(2-3), p.543-559.

Chung, H., Laforte, J., Reifschneider, D., and Williams, J.C., 2012. Have we underestimated the likelihood and severity of zero lower bound events?. *Journal of Money, Credit and Banking*, 44(1), p.47-82.

Cochrane, J.H., 1989. The return of the liquidity effect: A study of the short-run relation between money growth and interest rates. *Journal of Business and Economic Statistics*, 7(1), p.75-83.

Cochrane, J.H., and Piazzesi, M., 2002. The Fed and interest rates: A high-frequency identification. *The American Economic Review Papers and Proceedings*, 92(2), 90-95.

Cook, T., and Hahn, T., 1989. The effect of changes in the federal funds rate target on market interest rates in the 1970s. *Journal of Monetary Economics*, 24(3), p.331-351.

Cover, J.P., 1992. Asymmetric effects of positive and negative money-supply shocks. *The Quarterly Journal of Economics*, 107(4), p.1261-1282.

Culbertson, J., 1957. The term structure of interest rates. *The Quarterly Journal of Economics*, 71(4), p.485-517.

Curdia, V., and Woodford, M., 2011. The central-bank balance sheet as an instrument of monetary policy. *Journal of Monetary Economics*, 58(1), p.54-79.

D’Amico, S., and King, T.B., Flow and stock effects of large-scale treasury purchases: Evidence on the importance of local supply. *Journal of Financial Economics*, 108(2), p.425-448.

D'Amico, S., English, W., Lopez-Salido, D., and Nelson, E., 2012. The Federal Reserve's large-scale asset purchase programmes: rationale and effects. *Economic Journal*, 122(564), p.415-446.

Diebold, F.,X., and Li, C., 2006. Forecasting the term structure of government bond yields. *Journal of Econometrics*, 130(2), p.337-364.

Duffee, G.R., 2015. Expected inflation and other determinants of Treasury yields. Working paper (unpublished).

Edelberg, W., and Marshall, D.A., 1996. Monetary policy shocks and long-term interest rates. *Federal Reserve Bank of Chicago Economic Perspectives*, 20(2), 2-17.

Engsted, T., and Pedersen, T.Q., 2014. Housing market volatility in the OECD area: evidence from VAR based return decompositions. *Journal of Macroeconomics*, 42, p.91-103.

Engsted, T., and Tanggaard, C., 2001. The Danish stock and bond markets: comovement, return predictability and variance decomposition. *Journal of Empirical Finance*, 8(3), p.243–271.

Engsted, T., and Tanggaard, C., 2004. The comovement of US and UK stock markets. *European Financial Management*, 10(4), p.593–607.

Engsted, T., and Tanggaard, C., 2007. The comovement of U.S. and German bond markets. *International Review of Financial Analysis*, 16(2), p.172-182.

Engsted, T., Pedersen, T.Q., and Tanggaard, C., 2012. Pitfalls in VAR based return decompositions: a clarification. *Journal of Banking and Finance*, 36(5), p.1255-1265.

Estrella, A., and Mishkin, F.S., 1997. The predictive power of the term structure of interest rates in Europe and the United States: Implications for the European Central Bank. *European Economic Review*, 41(7), p.1375-1401.

Evans, C.L., and Kuttner, K.N., 1998. Can VARs describe monetary policy? *Federal Reserve Bank of Chicago Working Paper Series*, Working paper 98-19.

Evans, C.L., and Marshall, D.A., 1998. Monetary policy and the term structure of nominal interest rates: Evidence and theory. *Carnegie-Rochester Conference Series on Public Policy*, 49, p.53-111.

Fama, E.F, and Bliss, R., 1987. The information in long-maturity forward rates. *The American Economic Review*, 77(4), p. 680-692

Fama, E.F., 2013. Does the Fed control interest rates?. *The Review of Asset Pricing Studies*, 3(2), p.180-199.

Farka, M., and DaSilva, A., 2011. The Fed and the term structure: Addressing simultaneity within a structural VAR model. *Journal of Empirical Finance*, 18(5), p.935-952.

Faust, J., and Wright, J.H., 2008. Efficient forecast tests for conditional policy forecasts. *Journal of Econometrics*, 146(2), p.293-303.

Faust, J., Swanson, E.T., and Wright, J.H., 2004. Do Federal Reserve policy surprises reveal superior information about the economy? *Contributions to Macroeconomics*, 4(1), Article 10, p.1-31.

Fawley, B.W., and Neely, C.J., 2013. Four stories of quantitative easing. *Federal Reserve Bank of St. Louis Review*, 95(1), p.51-88.

Fawley, B.W., and Neely, C.J., 2014. The evolution of Federal Reserve policy and the impact of monetary policy surprises on asset prices. *Federal Reserve Bank of St. Louis Review*, 96(1), p.73-109.

Feldstein, M., and Chamberlain, G., 1973. Multimarket expectations and the rate of interest. *Journal of Money, Credit and Banking*, 5(4), p.873-902.

Fratzscher, M., Lo Duca, M., and Straub, R., 2013. On the international spillovers of US quantitative easing. *ECB Working Paper Series*, Working paper 1557.

Fricke, C., and Menkhoff, L., 2014. Financial conditions, macroeconomic factors and (un)expected bond excess returns. *Deutsche Bundesbank*, Discussion paper 35/2014.

Gagnon, J., Raskin, M., Remache, J., and Sack, B., 2011. The financial market effects of the Federal Reserve's large-scale asset purchases. *International Journal of Central Banking*, 7(1), p.3-43.

Gambacorta, L., Hofmann, B., and Peersman, G., 2014. The effectiveness of unconventional monetary policy at the zero lower bound: a cross-country analysis. *Journal of Money, Credit and Banking*, 46(4), p.615-642.

Garrett, I., and Priestley, R., 2012. Dividend growth, cash flow, and discount rate news. *Journal of Financial and Quantitative Analysis*, 47(5), p.1003-1028.

Gertler, M., and Karadi, P., 2013. QE 1 vs. 2 vs. 3. . . : A Framework for analysing large-scale asset purchases as a monetary policy tool. *International Journal of Central Banking*, 9(1), p.5-53.

Gertler, M., and Karadi, P., 2015. Monetary policy surprises, credit costs, and economic activity. *The American Economic Journal: Macroeconomics*, 7(1), p.44-76.

Gibson, W.E., 1970. Interest rates and monetary policy. *Journal of Political Economy*, 78(3), p.431-455.

Goyenko, R.Y., and Ukhov, A.D., 2009. Stock and bond market liquidity: A long-run empirical analysis. *Journal of Financial and Quantitative Analysis*, 44(1), p.189-212.

Greenwood, R., and Vayanos, D., 2014. Bond supply and excess bond returns. *The Review of Financial Studies*, 27(3), p.663-713.

Gurkaynak, R.S., Sack, B., and Swanson, E.T., 2005. Do actions speak louder than words? The response of asset prices to monetary policy actions and statements. *International Journal of Central Banking*, 1(1), p.55-93.

Gurkaynak, R.S., Sack, B., and Wright, J.H., 2007. The U.S. treasury yield curve: 1961 to the present. *Federal Reserve Board Finance and Economics Discussion Series*, Working paper 2006-28.

Gurkaynak, R.S., Sack, B.P., and Swanson, E.T., 2007. Market-based measures of monetary policy expectations. *Journal of Business and Economic Statistics*, 25(2), p.201-212.

Hamilton, J.D., 2008. Assessing monetary policy effects using daily federal funds futures contracts. *Federal Reserve Bank of St. Louis Review*, 90(4), p.377-393.

Hamilton, J.D., and Wu, J.C., 2012. The effectiveness of alternative monetary policy tools in a zero lower bound environment. *Journal of Money, Credit and Banking*, 44(1), p.3-46.

Harada, Y., and Masujima, M., 2009. Effectiveness and transmission mechanisms of Japan's quantitative monetary easing policy. *Japanese Economy*, 36(1), p.48-105.

Johnson, R.R., Buetow, G.W., Jensen, G.R., and Reilly, F.K., 2003. Monetary policy and fixed income returns. *The Quarterly Review of Economics and Finance*, 43(1), p. 133-146.

Joslin, S., Priebsch, M., and Singleton, K.J., 2014. Risk premiums in dynamic term structure models with unspanned macro risks. *The Journal of Finance*, 69(3), p.1197-1233.

Karras, G., 2013. Asymmetric effects of monetary policy with or without quantitative easing: empirical evidence for the US. *Journal of Economic Asymmetries*, 10(1), p.1-9.

Kimura, T., Kobayashi, H., Muranaga, J., and Ugai, H., 2003. The effect of the increase in the monetary base on Japan's economy at zero interest rates: an empirical analysis. *BIS Papers*, Working paper 19.

Kontonikas, A., and Kostakis, A., 2013. On monetary policy and stock market anomalies. *Journal of Business Finance and Accounting*, 40(7-8), p.1009-1042.

Kontonikas, A., MacDonald, R., and Saggu, A. (2013). Stock market reaction to Fed funds rate surprises: state dependence and the financial crisis. *Journal of Banking and Finance*, 37(11), p.4025-4037.

Krishnamurthy, A., and Vissing-Jorgensen, A., 2011. The effects of quantitative easing on interest rates: Channels and implications for policy. *NBER Working Paper Series*, Working paper 17555.

Krogstrup, S., Reynard, S., and Sutter, B., 2012. Liquidity effects of quantitative easing on long-term interest rates. *SNB Working Papers*, Working paper 2012-2.

Kuttner, K.N., 2001. Monetary policy surprises and interest rates: Evidence from the Fed funds futures market. *Journal of Monetary Economics*, 47(3), p.523-544.

Ludvigson, S.C., and Ng, S., 2009. Macro factors in bond risk premia. *The Review of Financial Studies*, 22(12), p.5027-5067.

Maio, P., 2014. Another look at the stock return response to monetary policy actions. *Review of Finance*, 18(1), p.321-371.

Martin, C., and Milas, C., 2012. Quantitative easing: A sceptical survey. *Oxford Review of Economic Policy*, 28(4), p.750-764.

Mishkin, F.S., 1981. Monetary policy and long-term interest rates: An efficient markets approach. *Journal of Monetary Economics*, 7(1), p.29-55.

Modigliani, F., and Sutch, R., 1966. Innovations in interest rate policy. *The American Economic Review*, 56(1-2), p.178-197.

Moench, E., 2012. Term structure surprises: The predictive content of curvature, level, and slope. *Journal of Applied Econometrics*, 27(4), p.574-602.

Moessner, R., 2013. Effects of explicit FOMC policy rate guidance on market interest rates. *DNB Working Papers*, Working paper 384.

Neely, C.J., 2015. Unconventional monetary policy had large international effects. *Journal of Banking and Finance*, 52, p.101-111.

Nelson, C.R., and Siegel, A.F., 1987. Parsimonious modelling of yield curves. *The Journal of Business*, 60(4), p.473-489.

Nimark, K., 2008. Monetary policy with signal extraction from the bond market. *Journal of Monetary Economics*, 55(8), p.1389-1400.

Nitschka, T., 2014. What news drive variation in Swiss and US bond and stock excess returns?. *Swiss Journal of Economics and Statistics*, 150(2), p.89-118.

Nozawa, Y., 2014. What drives the cross-section of credit spreads?: A variance decomposition approach. *Federal Reserve Board Finance and Economics Discussion Series*, Working paper 2014-62.

Piazzesi, M., 2005. Bond yields and the Federal Reserve. *Journal of Political Economy*, 113(2), p.311–344.

Piazzesi, M., and Schneider, M., 2007. Equilibrium yield curves. In: Acemoglu, D., Rogoff, K., and Woodford, M. eds. *NBER Macroeconomics Annual 2006*. Cambridge: The MIT Press, p.389-472.

Piazzesi, M., and Swanson, E.T., 2008. Futures prices as risk-adjusted forecasts of monetary policy. *Journal of Monetary Economics*, 55(4), p.677-691.

Poole, W., Rasche, R.H., and Thornton, D.L., 2002. Market anticipations of monetary policy actions. *Federal Reserve Bank of St. Louis Review*, 84(4), p. 65-94.

Raskin, M.D., 2013. The effects of the Federal Reserve's date-based forward guidance. *Federal Reserve Board Finance and Economics Discussion Series*, Working paper 2013-37.

Reichenstein, W., 1987. The impact of money on short-term interest rates. *Economic Inquiry*, 25(1), 67-82.

Rigobon, R., and Sack, B., 2004. The impact of monetary policy on asset prices. *Journal of Monetary Economics*, 51(8), p.1553-1575.

Rogers, J.H., Scotti, C., and Wright, J.H., 2014. Evaluating asset-market effects of unconventional monetary policy: a multi-country review. *Economic Policy*, 29(80), p.749-799.

Rolley, V.V., and Sellon, G.H.Jr., 1995. Monetary policy actions and long-term interest rates. *Federal Reserve Bank of Kansas City Economic Review*, 80(4), p.77-89.

Romer, C.D., and Romer, D.H., 2000. Federal Reserve information and the behaviour of interest rates. *The American Economic Review*, 90(3), p.429-457.

Romer, C.D., and Romer, D.H., 2004. A new measure of monetary shocks: Derivation and implications. *The American Economic Review*, 94(4), p.1055-1084.

Sarno, L., Thornton, D.L., and Valente, G., 2007. The empirical failure of the expectations hypothesis of the term structure of bond yields. *Journal of Financial and Quantitative Analysis*, 42(1), p.81-100.

Strongin, S., 1995. The identification of monetary policy disturbances. Explaining the liquidity puzzle. *Journal of Monetary Economics*, 35(3), p.463-497.

Thornton, D.L., 2005. Tests of the expectations hypothesis: Resolving the anomalies when the short-term rate is the Federal funds rate. *Journal of Banking and Finance*, 29(10), p.2541-2556.

Thornton, D.L., 2012. Evidence on the portfolio balance channel of quantitative easing. *Federal Reserve Bank of St. Louis Working Paper Series*, Working paper 2012-015A.

Thornton, D.L., 2014a. The identification of the response of interest rates to monetary policy actions using market-based measures of monetary policy shocks. *Oxford Economic Papers*, 66(1), p.67-87.

Thornton, D.L., 2014b. An evaluation of event-study evidence on the effectiveness of the FOMC's LSAP program: Are the announcement effects identified?. *Federal Reserve Bank of St. Louis Working Paper Series*, Working paper 2013-033B.

Valckx, N., 2004. The decomposition of US and Euro area stock and bond returns and their sensitivity to economic state variables. *The European Journal of Finance*, 10(2), p.149-173.

Vayanos, D., and Vila, J-L., 2009. A preferred-habitat model of the term structure of interest rates. *NBER Working Paper Series*, Working paper 15487.

Vazquez, J., Maria-Dolores, R., and Londono, J., 2013. On the informational role of term structure in the US monetary policy rule. *Journal of Economic Dynamics and Control*, 37(9), p.1852-1871.

Vuolteenaho, T., 2002. What drives firm-level stock returns? *The Journal of Finance*, 57(1), p.233-264.

Wachter, J.A., 2006. A consumption-based model of the term structure of interest rates. *Journal of Financial Economics*, 79(2), p.365-399.

Walsh, C.E., 2003. *Monetary theory and policy*, 2nd ed. Cambridge: MIT Press.

Wieland, J.F., and Yang, M., 2015. Financial dampening. Working paper (unpublished).

Wright, J.H., 2012. What does monetary policy do to long-term interest rates at the zero lower bound?. *The Economic Journal*, 122(564), p.F447-F466.

Wu, T., 2014. Unconventional monetary policy and long-term interest rates. *IMF Working Papers*, Working paper WP/14/189.

Chapter 2 - Appendix A

This Appendix provides the summary of the derivation of the log-linear relationship between current unexpected excess bond returns, expected future excess returns, inflation and real interest rates. The derivation is in line with Campbell and Ammer (1993).

The gross nominal holding-period return $\left(1+R_{n,t+1}\right)$ on an n-period bond from t to t+1 is:

$$(1+R_{n,t+1})=\frac{P_{n-1,t+1}}{P_{n,t}}=\frac{(1+Y_{n,t})^{n}}{(1+Y_{n-1,t+1})^{n-1}} \tag{A2.1}$$

where $P_{n,t}$ and $Y_{n,t}$ denote the price and yield on an n-period zero-coupon bond at time t. Taking logs on both sides of Equation (A2.1), the log nominal holding-period return is obtained:

$$r_{n,t+1} = p_{n-1,t+1} - p_{n,t} = y_{n,t} - (n-1)(y_{n-1,t+1} - y_{n,t}) \tag{A2.2}$$

Re-arranging (A2.2) in terms of the current log bond price and solving forward:

$$p_{n,t} = -\sum_{j=0}^{n-1} r_{n-j,t+1+j} \tag{A2.3}$$

Taking expectations at time t on both sides of Equation (A2.3) it can be written as follows:

$$p_{n,t} = -E_t\left[\sum_{j=0}^{n-1} r_{n-j,t+1+j}\right] \tag{A2.4}$$

Using Equations (A2.4) and (A2.2) one can obtain an expression for current unexpected bond returns which shows that they are negatively related to the revisions in expectations about future bond returns:

$$r_{n,t+1} - E_t\left[r_{n,t+1}\right] = -(E_{t+1}-E_t)\left[\sum_{j=1}^{n-1} r_{n-j,t+1+j}\right] \tag{A2.5}$$

Following Campbell, Lo and MacKinlay (1997, p.414), excess bond returns are defined as follows:

$$x_{n,t+1} = r_{n,t+1} - y_{1,t} = r_{n,t+1} - \pi_{t+1} - r^{i}_{t+1} \tag{A2.6}$$

where $y_{1,t}$ is the log nominal short-term risk-free rate at time t, π_{t+1} is the inflation rate between t and t+1 (defined as the log difference of the consumer price index), and r^{i}_{t+1} is the real interest rate at time t+1.

Using Equation (A2.6), (A2.5) can be re-written in terms of excess bond returns and then one can obtain (A2.7) which corresponds to Equation (2.1) in the main text:

$$\begin{aligned}\tilde{x}_{n,t+1} &= x_{n,t+1} - E_t\left[x_{n,t+1}\right] = (E_{t+1} - E_t)\left[-\sum_{j=1}^{n-1} x_{n-j,t+1+j} - \sum_{j=1}^{n-1} \pi_{t+1+j} - \sum_{j=1}^{n-1} r^{i}_{t+1+j}\right] \\ &= -\tilde{x}_{x,t+1} - \tilde{x}_{\pi,t+1} - \tilde{x}_{r^{i},t+1}\end{aligned} \tag{A2.7}$$

Chapter 2 – Appendix B

This Appendix shows how empirical proxies for the revisions in expectations in Equation (2.1) can be obtained using the VAR approach. The analysis is based upon Campbell and Ammer (1993). The starting point is a first order VAR model:

$$Z_{t+1} = AZ_t + W_{t+1} \tag{B2.1}$$

where Z_t is the vector of endogenous state variables, A denotes the matrix of VAR parameters, and W_t is the vector of forecast residuals.

The state vector contains the change in nominal short-term risk-free rate ($\Delta y_{1,t}$), the spread between long-term and short-term yields ($s_{n,t}$), the real interest rate (r_t^i) and the relative bill rate rb_t, i.e. the difference between the short-term nominal interest rate and its 12-month backwards moving average: $rb_t = y_{1,t} - \left(\frac{1}{12}\right)\sum_{i=1}^{12} y_{1,t-i}$.

Innovations to one-period excess bond returns at time t+1 ($\tilde{x}_{n,t+1}$) are related to the innovations in the nominal short-term risk-free rate ($\tilde{y}_{1,t+1}$) and innovations in the yield spread between (n-1)-period and 1-period bonds ($\tilde{s}_{n-1,t+1}$):

$$\tilde{x}_{n,t+1} = -(n-1)\,\tilde{y}_{n-1,t+1} = -(n-1)(\tilde{y}_{1,t+1} + \tilde{s}_{n-1,t+1}) \tag{B2.2}$$

Hence, the first and second equations in the VAR model are used to extract the proxy for unexpected excess bond returns at time t+1:

$$\tilde{x}_{n,t+1} = -(n-1)(s_1^T W_{t+1} + s_2^T W_{t+1}) \tag{B2.3}$$

where s_i^T is the unit selection vector with i representing the i^{th} equation in the VAR model and accordingly the i^{th} element of the vector is set to 1. For instance, s_1^T is the vector that takes the value of one in the cell corresponding to the position of the first variable in the VAR ($\Delta y_{1,t}$).

This approach is appropriate since innovations in the level of the nominal short-term risk-free rate are the same as innovations to the change in the short rate, given that the lagged rate is known to the investors beforehand. Furthermore, the distinction between $\tilde{s}_{n-1,t+1}$ and $\tilde{s}_{n,t+1}$ can be safely ignored given that the approximation error becomes very small as n increases.

To obtain the estimates of the revisions in expectations about future real interest rates, one can use the projections from the error vector:

$$(E_{t+1} - E_t)\left[Z_{t+1+j}\right] = A^j W_{t+1} \tag{B2.4}$$

Real interest rates news is estimated using information the third equation in the VAR:

$$\tilde{x}_{r^i,t+1} = s_3^T \sum_{j=1}^{n-1} A^j W_{t+1} \tag{B2.5}$$

Using the geometric series properties it can be shown that Equation (B2.5) becomes:

$$\tilde{x}_{r^i,t+1} = s_3^T \left(\frac{A^1 - A^{n-1+1}}{I - A} \right) W_{t+1} = s_3^T (I - A)^{-1} (A - A^n) W_{t+1} \tag{B2.6}$$

where I is the identity matrix.

Inflation news is calculated using information about nominal short-term interest rates and real interest rates:

$$\begin{aligned} \tilde{x}_{\pi,t+1} &= (E_{t+1} - E_t) \sum_{j=1}^{n-1} \pi_{t+1+j} = (E_{t+1} - E_t) \sum_{j=1}^{n-1} (y_{1,t+j} - r^i_{t+1+j}) \\ &= (E_{t+1} - E_t) \sum_{j=1}^{n-1} y_{1,t+j} \ - \tilde{x}_{r^i,t+1} \end{aligned} \tag{B2.7}$$

Since the VAR state vector contains the first difference of the nominal short-term interest rate, the first term in (B2.7) is converted to the weighted sum of the first differences of the short rate:

$$\tilde{x}_{\pi,t+1} = (E_{t+1} - E_t) \sum_{j=1}^{n-1} (n - j) \Delta y_{1,t+j} \ - \tilde{x}_{r^i,t+1} \tag{B2.8}$$

It can be shown that Equation (B2.8) can be re-written as follows:

$$\tilde{x}_{\pi,t+1} = s_1^T \left\{ \left(I - A\right)^{-1} \left[\left(n-1\right) I + \left(I - A\right)^{-1} \left(A^n - A\right) \right] \right\} W_{t+1} - \tilde{x}_{r^i,t+1} \tag{B2.9}$$

Finally, the estimates for revisions in future excess bond returns are obtained as the residual component using Equation (2.1):

$$\tilde{x}_{x,t+1} = -\tilde{x}_{n,t+1} - \tilde{x}_{r^i,t+1} - \tilde{x}_{\pi,t+1} \tag{B2.10}$$

Chapter 2 – Appendix C

Table C2.1: Descriptive statistics and unit root tests

Variables	1985:1–2014:2				1985:1–2007:7			
Panel A	Mean	St. dev.	Min.	Max.	Mean	St. dev.	Min.	Max.
$\Delta y_{1,t}$	-0.0219	0.5397	-2.9858	2.0324	-0.0106	0.5971	-2.9858	2.0324
$s_{120,t}$	2.1306	1.3374	-1.0859	5.0091	1.9888	1.4003	-1.0859	5.0091
$s_{60,t}$	1.4257	1.0714	-1.5288	4.4923	1.4530	1.1549	-1.5288	4.4923
$s_{24,t}$	0.7751	0.8091	-1.7930	3.6581	0.8926	0.8636	-1.7930	3.6581
r_i^i	0.9893	3.5285	-12.8149	22.3965	1.6556	2.9451	-12.8150	12.9246
rb_t	-0.1760	0.8809	-2.7729	2.4335	-0.1089	0.9239	-2.4029	2.4335
ΔFFR_t	-0.0237	0.2066	-0.9600	0.8700	-0.0115	0.2136	-0.6600	0.8700
ΔFFR_t^U	-0.0313	0.0933	-0.6265	0.3300	-0.0414	0.1017	-0.6265	0.3125
ΔMB_t	0.8570	2.3134	-8.4381	25.9621	0.5484	0.5657	-2.6733	3.8956
ΔMB_t^U	-0.0000	1.6363	-9.6332	12.7958	-0.1033	0.7291	-4.3168	2.5903

Variables	ADF constant	ADF constant & trend	PP constant	PP constant & trend
Panel B				
$\Delta y_{1,t}$	-3.69 [11]***	-3.68 [11]**	-27.39 [7]***	-27.36 [7]***
$s_{120,t}$	-4.08 [12]***	-4.08 [12]***	-4.28 [8]***	-4.27 [8] ***
$s_{60,t}$	-4.15 [12]***	-4.30 [12]***	-5.57 [7]***	-5.90 [8]***
$s_{24,t}$	-4.09 [12]***	-5.04 [12]***	-8.07 [9]***	-9.19 [9]***
r_i^i	-2.81 [14]*	-3.64 [14]**	-9.94 [4]***	-11.12 [4]***
rb_t	-4.46 [15]***	-4.45 [15]***	-6.38 [9]***	-6.38 [9]***
ΔFFR_t	-5.01 [4]***	-5.00 [4]***	-11.64 [10]***	-11.63 [10]***
ΔFFR_t^U	-2.58 [11]*	-4.68 [14]***	-14.65 [5]***	-15.62 [7]***
ΔMB_t	-5.43 [8]***	-7.59 [6]***	-8.85 [5]***	-9.00 [6]***
ΔMB_t^U	-18.59 [0]***	-18.70 [0]***	-18.59 [2]***	-18.70 [0]***

Notes: Panel A of this table reports the summary statistics for variables used for the benchmark VAR estimations as well as four indicators of monetary policy actions over the full sample period (1985:1 – 2014:2) and pre-crisis period (1985:1 – 2007:7); the first difference of the 1-month Treasury bill rate (Δy_1), the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill (s_{120}, s_{60} and s_{24}, respectively), the real interest rate (r^i) and the relative bill rate (*rb*); the change in the federal funds rate (ΔFFR), the unexpected FFR change (ΔFFR^U), the change in log monetary base (ΔMB) and the unexpected change in log monetary base (ΔMB^U). Due to data availability on FFR futures, in the case of the unexpected change in the FFR, the full sample commences in 1989:2. Panel B of this table reports the full sample test statistics for the augmented Dickey-Fuller (ADF) and Phillips Perron (PP) unit root tests with (a) constant and (b) constant and trend. In brackets the lag-length of the ADF test is reported, based on Akaike information criterion, and the Newey-West bandwidth for the PP test. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.2: Fed announcements and balance sheet developments

Date	Facility/programme	Description	Source
2008:9-11	Liquidity facilities: Balance sheet expansion	Increased usage of the existing and newly setup liquidity facilities led to a substantial increase in the Fed's balance sheet as operations were no longer sterilised.	Federal Reserve: Recent balance sheet trends
2008:11	QE1 announced: Further balance sheet expansion	Federal Reserve announces purchases of up to $100 billion in direct obligations of housing-related government-sponsored enterprises (GSEs) and of up to $500 billion in agency mortgage-backed securities (MBS).	FOMC statement
2008:12	QE1 expansion hint	First hint on purchases of Treasuries: "…the Fed could purchase longer-term Treasury…in substantial quantities".	Chairman Bernanke's speech
2008:12	QE1 expansion hint	FOMC considers QE extension to Treasuries: "The Committee is also evaluating the potential benefits of purchasing longer-term Treasury securities".	FOMC statement
2009:1	QE1 expansion hint	FOMC confirms the intention to purchase Treasuries: "The Committee also is prepared to purchase longer-term Treasury securities".	FOMC statement
2009:3	QE1 extended: Further balance sheet expansion	FOMC announces additional purchases of $750 billion in MBS, $100 billion in GSE debt and of up to $300 billion in longer-term Treasuries over the next six months.	FOMC statement
2010:8	QE2 hint	Chairman Bernanke hints about QE2: "…the Committee is prepared to provide additional monetary accommodation through unconventional measures".	Chairman Bernanke's speech
2010:11	QE2 announced: Further balance sheet expansion	FOMC announces additional purchases of $600 billion in Treasuries ($75 billion per month) by the end of the second quarter of next year.	FOMC statement
2011:9	Operation Twist	FOMC announces purchases of $400 billion in Treasuries with remaining maturities of 6 to 30 years and $400 billion sales of Treasuries maturing in 3 or less years.	FOMC statement
2012:6	Operation Twist extension	Programme extended through to the end of 2012.	FOMC statement
2012:8	QE3 hint	FOMC considers additional stimulus: "…additional monetary accommodation would likely be warranted fairly soon."	FOMC minutes
2012:9	QE3announced: Further balance sheet expansion	FOMC announces additional purchases of MBS ($40 billion per month).	FOMC statement
2012:12	QE3 extended: Further balance sheet expansion	FOMC announces additional purchases of longer-term Treasuries ($45 billion per month).	FOMC statement

Notes: This table reports the months that were associated with the Federal Reserve announcements and policymakers' speeches related to unconventional policies, provides details about their content and lists the sources. The liquidity facilities include: central banks liquidity swaps, Primary Dealer Credit Facility, Asset-Backed Commercial Paper Money Market Mutual Fund Liquidity Facility, primary and secondary credit, seasonal credit, Commercial Paper Funding Facility, and Term Auction Facility. More details are provided by the Federal Reserve at http://www.federalreserve.gov/monetarypolicy/bst_recenttrends.htm.

Table C2.3: Impact of macroeconomic news on monetary policy indicators

Macroeconomic surprise	1991:10 – 2014:2			
	ΔFFR	ΔFFR^U	ΔMB	ΔMB^U
CPI inflation	0.064 (0.087)	-0.050 (0.041)	-2.883 (2.448)	-2.160 (1.761)
Core CPI inflation	0.002 (0.148)	0.038 (0.046)	-0.030 (2.092)	1.190 (1.581)
Nonfarm payrolls	0.007 (0.012)	-0.005 (0.005)	-0.083 (0.161)	-0.050 (0.095)
Industrial production	0.020 (0.031)	-0.020* (0.011)	-0.286 (0.448)	-0.076 (0.419)
Retail sales excl. autos	0.002 (0.018)	0.009 (0.009)	0.286 (0.494)	0.437 (0.511)
R^2	0.006	0.020	0.026	0.032
	1991:10 – 2007:7			
CPI inflation	-0.002 (0.091)	-0.064 (0.044)	0.337 (0.452)	0.991* (0.599)
Core CPI inflation	-0.156 (0.169)	0.090 (0.057)	-0.218 (0.527)	-0.344 (0.741)
Nonfarm payrolls	0.002 (0.014)	-0.005 (0.006)	0.019 (0.038)	0.009 (0.048)
Industrial production	0.037 (0.045)	-0.010 (0.018)	-0.158 (0.106)	-0.220 (0.180)
Retail sales excl. autos	-0.015 (0.029)	0.022 (0.016)	-0.112 (0.200)	-0.014 (0.321)
R^2	0.012	0.024	0.014	0.035

Notes: This table reports the estimated parameters from the regressions of monetary policy indicators on macroeconomic surprises. The monetary indicators are the change in the federal funds rate (ΔFFR), the unexpected FFR change (ΔFFR^U), the change in log monetary base (ΔMB) and the unexpected change in log monetary base (ΔMB^U). The macroeconomic surprises relate to Reuters Economic Polls and are calculated based on 'Actual' (the actual value that was reported by the primary source) minus 'Median Forecast' (the forecast figure from the polls prior to the announcement) after the actual value is released. The following macroeconomic variables are considered: CPI inflation, core CPI inflation, change in nonfarm payrolls, growth rate of industrial production and growth rate of retail sales (excluding autos). The upper panel of the table provides the full sample (1991:10 – 2014:2) estimates, while the pre-crisis period (1991:10 – 2007:7) estimates are shown in the lower panel. Due to data availability on macroeconomic surprises, the full sample commences in 1991:10. Data is obtained from the Datastream. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.4: Impact of monetary policy on excess bond returns (since October 1992) – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1992:10 – 2014:2**	**1992:10 – 2007:7**	**1992:10 – 2014:2**	**1992:10 – 2007:7**	**1992:10 – 2014:2**	**1992:10 – 2007:7**
$\tilde{x}_n^{MP}$	-9.196	-1.606	-10.978**	-3.907	-6.727***	-4.053**
	(6.307)	(6.466)	(4.108)	(4.249)	(1.839)	(1.962)
$\tilde{x}_{r^i}^{MP}$	1.805	3.519	-1.450	0.306	-2.470**	-0.940*
	(2.192)	(2.153)	(1.216)	(1.138)	(1.192)	(0.560)
$\tilde{x}_\pi^{MP}$	28.431***	15.705	18.768***	8.185	11.571***	6.164***
	(10.347)	(10.358)	(4.527)	(5.224)	(1.663)	(1.502)
$\tilde{x}_x^{MP}$	-21.040**	-17.617**	-6.340*	-4.584*	-2.374**	-1.164*
	(8.812)	(7.550)	(3.306)	(2.571)	(1.029)	(0.628)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the alternative full sample (1992:10 – 2014:2) and pre-crisis period (1992:10 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.5: Impact of monetary policy on excess bond returns (since October 1992) – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1992:10 – 2014:2**	**1992:10 – 2007:7**	**1992:10 – 2014:2**	**1992:10 – 2007:7**	**1992:10 – 2014:2**	**1992:10 – 2007:7**
$\tilde{x}_n^{MP}$	36.748***	-15.600***	-0.443	-13.134***	-11.626***	-13.511***
	(2.500)	(2.536)	(1.427)	(1.246)	(0.885)	(0.540)
$\tilde{x}_{r^i}^{MP}$	-0.687	-3.356**	0.263	-1.330	3.068	1.511
	(1.453)	(1.366)	(0.935)	(1.163)	(2.589)	(2.724)
$\tilde{x}_\pi^{MP}$	-45.644***	3.273	-5.377	9.446**	7.581*	11.456***
	(5.850)	(6.324)	(3.312)	(3.956)	(4.038)	(3.914)
$\tilde{x}_x^{MP}$	9.583	15.684***	5.557**	5.018	0.976	0.544
	(6.670)	(5.574)	(2.370)	(3.274)	(2.046)	(2.230)

Notes: This table reports the impact of an unexpected change in the Federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.4 notes.

Table C2.6: Impact of monetary policy on excess bond returns (since February 1994) – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1994:2 – 2014:2**	**1994:2 – 2007:7**	**1994:2 – 2014:2**	**1994:2 – 2007:7**	**1994:2 – 2014:2**	**1994:2 – 2007:7**
$\tilde{x}_n^{MP}$	-9.991 (6.353)	-2.411 (6.610)	-11.709*** (4.142)	-4.872 (4.349)	-6.900*** (1.856)	-4.277** (2.001)
$\tilde{x}_{r^i}^{MP}$	1.727 (2.229)	3.393 (2.193)	-1.571 (1.254)	0.112 (1.134)	-2.600** (1.251)	-1.133* (0.627)
$\tilde{x}_\pi^{MP}$	29.684*** (10.515)	16.905 (10.580)	19.783*** (4.566)	9.447* (5.193)	11.967*** (1.729)	6.695*** (1.520)
$\tilde{x}_x^{MP}$	-21.420** (9.009)	-17.886** (7.712)	-6.503* (3.423)	-4.687* (2.618)	-2.468** (1.067)	-1.284* (0.675)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the alternative full sample (1994:2 – 2014:2) and pre-crisis period (1994:2 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.7: Impact of monetary policy on excess bond returns (since February 1994) – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1994:2 – 2014:2**	**1994:2 – 2007:7**	**1994:2 – 2014:2**	**1994:2 – 2007:7**	**1994:2 – 2014:2**	**1994:2 – 2007:7**
$\tilde{x}_n^{MP}$	54.330*** (3.861)	-0.046 (3.252)	10.088*** (1.283)	-0.096 (1.502)	-8.470*** (0.692)	-9.028*** (0.455)
$\tilde{x}_{r^i}^{MP}$	1.630 (1.880)	-1.153 (1.105)	1.686 (1.039)	-0.061 (0.845)	-3.862* (1.993)	1.884 (1.839)
$\tilde{x}_\pi^{MP}$	-62.918*** (7.706)	-11.526** (5.791)	-17.733*** (2.692)	-4.731 (2.894)	3.005 (3.136)	6.057** (2.674)
$\tilde{x}_x^{MP}$	6.958 (9.085)	12.725** (4.969)	5.959** (2.863)	4.887** (2.170)	1.603 (1.635)	1.087 (1.589)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.6 notes.

Table C2.8: Variance decomposition for excess bond returns – alternative VAR specification [1] – adding industrial production growth

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var\left(\tilde{x}_{\pi}\right)$	0.832*** (0.264)	0.814*** (0.246)	1.111** (0.433)	1.128** (0.444)	1.636** (0.787)	1.708** (0.697)
$2Cov\left(\tilde{x}_{\pi}, \tilde{x}_{r^i}\right)$	-0.086 (0.083)	-0.085 (0.086)	-0.190 (0.192)	-0.226 (0.186)	-0.698 (0.717)	-0.685 (0.582)
$2Cov\left(\tilde{x}_{\pi}, \tilde{x}_{x}\right)$	0.048 (0.284)	0.113 (0.216)	-0.102 (0.356)	-0.017 (0.338)	-0.592 (0.472)	-0.377 (0.436)
$Var\left(\tilde{x}_{r^i}\right)$	0.018* (0.010)	0.017 (0.012)	0.039* (0.023)	0.035 (0.024)	0.272* (0.148)	0.145 (0.107)
$2Cov\left(\tilde{x}_{r^i}, \tilde{x}_{x}\right)$	-0.068 (0.056)	-0.103* (0.060)	0.026 (0.068)	-0.050 (0.069)	0.269 (0.180)	0.102 (0.129)
$Var\left(\tilde{x}_{x}\right)$	0.256 (0.167)	0.244 (0.184)	0.117 (0.101)	0.131 (0.115)	0.113 (0.079)	0.107 (0.074)
$R^2\left(\tilde{x}_{\pi}\right)$	0.795*** (0.153)	0.842*** (0.147)	0.837*** (0.114)	0.897*** (0.089)	0.600** (0.157)	0.811*** (0.102)
$R^2\left(\tilde{x}_{r^i}\right)$	0.197 (0.145)	0.338** (0.151)	0.049 (0.132)	0.310 (0.197)	0.012 (0.089)	0.147 (0.239)
$R^2\left(\tilde{x}_{x}\right)$	0.236 (0.203)	0.254 (0.197)	0.054 (0.219)	0.073 (0.243)	0.021 (0.134)	0.009 (0.091)

Notes: This table reports the variance decomposition of unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_x$) and the covariances between these three components. News components are extracted from a VAR(1) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the first difference in log industrial production index and the relative bill rate. The first and second column for each bond maturity report the full sample (1985:1 – 2014:2) and pre-crisis period (1985:1 – 2007:7) results, respectively. R^2 values are obtained from regressions of unexpected excess returns on each news component. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.9: Variance decomposition for excess bond returns – alternative VAR specification [2] – adding unemployment rate

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var\left(\tilde{x}_{\pi}\right)$	0.838** (0.397)	1.578 (0.978)	0.833** (0.361)	1.703* (0.894)	1.269** (0.610)	1.797** (0.761)
$2Cov\left(\tilde{x}_{\pi}, \tilde{x}_{r^i}\right)$	0.227 (0.183)	-0.093 (0.180)	0.199 (0.174)	-0.186 (0.302)	-0.357 (0.472)	-0.632 (0.625)
$2Cov\left(\tilde{x}_{\pi}, \tilde{x}_{x}\right)$	-0.264 (0.510)	-0.688 (1.206)	-0.280 (0.385)	-0.709 (0.963)	-0.509 (0.386)	-0.495 (0.503)
$Var\left(\tilde{x}_{r^i}\right)$	0.063 (0.051)	0.012 (0.008)	0.133* (0.080)	0.026 (0.019)	0.303** (0.122)	0.134 (0.105)
$2Cov\left(\tilde{x}_{r^i}, \tilde{x}_{x}\right)$	-0.061 (0.148)	-0.026 (0.076)	0.004 (0.129)	0.006 (0.104)	0.193 (0.151)	0.106 (0.140)
$Var\left(\tilde{x}_{x}\right)$	0.197 (0.159)	0.217 (0.303)	0.110 (0.096)	0.160 (0.196)	0.102 (0.075)	0.090 (0.062)
$R^2\left(\tilde{x}_{\pi}\right)$	0.801*** (0.152)	0.894*** (0.107)	0.755*** (0.152)	0.926*** (0.067)	0.550*** (0.175)	0.846*** (0.088)
$R^2\left(\tilde{x}_{r^i}\right)$	0.339** (0.150)	0.193 (0.327)	0.415** (0.188)	0.162 (0.362)	0.161 (0.271)	0.124 (0.249)
$R^2\left(\tilde{x}_{x}\right)$	0.006 (0.084)	0.090 (0.304)	0.007 (0.103)	0.229 (0.449)	0.031 (0.170)	0.121 (0.349)

Notes: This table reports the variance decomposition of unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_{x}$) and the covariances between these three components. News components are extracted from a VAR(1) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the civilian unemployment rate and the relative bill rate. See also Table C2.8 notes.

Table C2.10: Variance decomposition for excess bond returns – alternative VAR specification [3] – adding Chicago Fed National Activity Index

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var(\tilde{x}_{\pi})$	0.862*** (0.298)	0.847*** (0.253)	1.146** (0.467)	1.178** (0.456)	1.655** (0.825)	1.781** (0.722)
$2Cov(\tilde{x}_{\pi}, \tilde{x}_{r^i})$	-0.082 (0.085)	-0.085 (0.089)	-0.188 (0.203)	-0.230 (0.195)	-0.693 (0.744)	-0.708 (0.615)
$2Cov(\tilde{x}_{\pi}, \tilde{x}_{x})$	-0.003 (0.350)	0.046 (0.230)	-0.171 (0.413)	-0.111 (0.365)	-0.643 (0.528)	-0.496 (0.465)
$Var(\tilde{x}_{r^i})$	0.018* (0.010)	0.018 (0.012)	0.038 (0.023)	0.036 (0.025)	0.268* (0.151)	0.150 (0.114)
$2Cov(\tilde{x}_{r^i}, \tilde{x}_{x})$	-0.071 (0.057)	-0.110* (0.061)	0.022 (0.075)	-0.056 (0.072)	0.260 (0.206)	0.107 (0.140)
$Var(\tilde{x}_{x})$	0.277 (0.184)	0.283 (0.185)	0.152 (0.114)	0.183 (0.117)	0.154 (0.097)	0.168** (0.081)
$R^2(\tilde{x}_{\pi})$	0.778*** (0.167)	0.809*** (0.145)	0.816*** (0.122)	0.861*** (0.089)	0.588*** (0.153)	0.780*** (0.095)
$R^2(\tilde{x}_{r^i})$	0.195 (0.152)	0.347** (0.153)	0.051 (0.137)	0.320 (0.198)	0.010 (0.078)	0.153 (0.233)
$R^2(\tilde{x}_{x})$	0.208 (0.208)	0.223 (0.184)	0.039 (0.174)	0.054 (0.186)	0.009 (0.075)	0.004 (0.050)

Notes: This table reports the variance decomposition of the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_{x}$) and the covariances between these three components. News components are extracted from a VAR(1) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the Chicago Fed National Activity Index and the relative bill rate. See also Table C2.8 notes.

Table C2.11: Variance decomposition for excess bond returns – alternative VAR specification [4] – adding Chicago Fed Adjusted National Financial Conditions Index

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	1985:1–2007:7	1985:1–2014:2	1985:1–2007:7	1985:1–2014:2	1985:1–2007:7
$Var\left(\tilde{x}_{\pi}\right)$	0.847** (0.341)	0.757*** (0.218)	1.181** (0.568)	1.059** (0.414)	1.911* (1.118)	1.860** (0.814)
$2Cov\left(\tilde{x}_{\pi},\tilde{x}_{r^i}\right)$	-0.124 (0.136)	-0.085 (0.092)	-0.283 (0.323)	-0.241 (0.212)	-1.106 (1.201)	-0.990 (0.782)
$2Cov\left(\tilde{x}_{\pi},\tilde{x}_{x}\right)$	-0.079 (0.470)	0.064 (0.270)	-0.280 (0.559)	-0.082 (0.359)	-0.876 (0.717)	-0.538 (0.502)
$Var\left(\tilde{x}_{r^i}\right)$	0.025 (0.017)	0.023 (0.015)	0.062 (0.052)	0.054 (0.038)	0.419 (0.320)	0.278 (0.198)
$2Cov\left(\tilde{x}_{r^i},\tilde{x}_{x}\right)$	-0.019 (0.099)	-0.052 (0.079)	0.115 (0.158)	0.031 (0.116)	0.471 (0.366)	0.254 (0.220)
$Var\left(\tilde{x}_{x}\right)$	0.351 (0.241)	0.293 (0.218)	0.206 (0.174)	0.178 (0.166)	0.182 (0.134)	0.135 (0.104)
$R^2\left(\tilde{x}_{\pi}\right)$	0.656*** (0.248)	0.737*** (0.218)	0.685** (0.207)	0.761*** (0.198)	0.443** (0.185)	0.646*** (0.180)
$R^2\left(\tilde{x}_{r^i}\right)$	0.090 (0.118)	0.092 (0.145)	0.008 (0.049)	0.047 (0.129)	0.025 (0.109)	0.029 (0.112)
$R^2\left(\tilde{x}_{x}\right)$	0.260 (0.197)	0.305* (0.171)	0.074 (0.204)	0.131 (0.251)	0.002 (0.035)	0.000 (0.016)

Notes: This table reports the variance decomposition of the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_{x}$) and the covariances between these three components. News components are extracted from a VAR(1) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the Chicago Fed Adjusted National Financial Conditions Index and the relative bill rate. See also Table C2.8 notes.

Table C2.12: Impact of monetary policy on excess bond returns with alternative VAR specification [1] – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-22.553***	-21.550***	-15.517***	-13.798***	-8.235***	-7.734***
	(5.369)	(4.810)	(3.044)	(2.831)	(1.379)	(1.317)
$\tilde{x}_{r^i}^{MP}$	0.845	1.494	-1.518	-0.716	-2.071	-1.264
	(1.785)	(1.635)	(1.245)	(1.071)	(1.545)	(1.338)
$\tilde{x}_\pi^{MP}$	36.576***	33.075***	22.921***	19.602***	12.847***	11.008***
	(8.184)	(7.433)	(3.672)	(3.333)	(1.886)	(1.617)
$\tilde{x}_x^{MP}$	-14.868*	-13.019*	-5.887*	-5.088	-2.541**	-2.011*
	(8.034)	(6.977)	(3.555)	(3.093)	(1.247)	(1.117)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the first difference of log industrial production index and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.13: Impact of monetary policy on excess bond returns with alternative VAR specification [1] – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-26.994***	-56.238***	-25.408***	-34.603***	-16.565***	-18.633***
	(4.168)	(4.880)	(2.097)	(2.203)	(0.973)	(0.887)
$\tilde{x}_{r^i}^{MP}$	-2.163*	-2.518*	-0.469	-0.736	2.251	1.934
	(1.195)	(1.512)	(1.613)	(1.936)	(3.032)	(3.349)
$\tilde{x}_\pi^{MP}$	17.993***	44.666***	23.389***	33.236***	15.067***	17.981***
	(5.609)	(6.209)	(5.553)	(5.785)	(4.420)	(4.536)
$\tilde{x}_x^{MP}$	11.164**	14.090*	2.487	2.103	-0.753	-1.282
	(5.000)	(7.587)	(4.541)	(5.732)	(2.396)	(2.651)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.12 notes.

Table C2.14: Impact of monetary policy on excess bond returns with alternative VAR specification [1] – MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.916*** (0.337)	-1.078 (1.055)	0.790*** (0.114)	0.757 (0.538)	0.313*** (0.041)	0.531** (0.212)
$\tilde{x}_{r^i}^{MP}$	0.166* (0.100)	-0.538** (0.267)	0.260** (0.090)	-0.477** (0.211)	0.282*** (0.095)	-0.483** (0.205)
$\tilde{x}_\pi^{MP}$	-2.128*** (0.360)	0.668 (1.530)	-1.494*** (0.205)	-0.214 (0.783)	-0.799*** (0.122)	0.126 (0.396)
$\tilde{x}_x^{MP}$	1.046* (0.598)	0.948 (0.858)	0.445* (0.239)	-0.066 (0.383)	0.205*** (0.073)	-0.174 (0.151)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.12 notes.

Table C2.15: Impact of monetary policy on excess bond returns with alternative VAR specification [1] – Unexpected MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	1.202*** (0.198)	1.159** (0.557)	1.179*** (0.087)	1.069*** (0.329)	0.478*** (0.034)	0.339** (0.146)
$\tilde{x}_{r^i}^{MP}$	0.306** (0.123)	0.024 (0.253)	0.342** (0.138)	-0.300 (0.205)	0.308* (0.157)	-0.532*** (0.177)
$\tilde{x}_\pi^{MP}$	-2.174** (0.588)	1.622 (1.235)	-1.858*** (0.406)	0.403 (1.716)	-0.995*** (0.229)	0.563* (0.314)
$\tilde{x}_x^{MP}$	0.666 (0.571)	-2.805** (1.125)	0.336 (0.300)	-1.172** (0.518)	0.210** (0.104)	-0.371** (0.153)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.12 notes.

Table C2.16: Impact of monetary policy on excess bond returns with alternative VAR specification [2] – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-21.101*** (5.042)	-20.791*** (5.145)	-15.450*** (3.014)	-13.603*** (2.919)	-8.508*** (1.351)	-7.832*** (1.293)
$\tilde{x}_{r^i}^{MP}$	5.369* (3.156)	5.387* (2.794)	2.761 (2.076)	2.750 (1.782)	-0.080 (1.682)	0.285 (1.473)
$\tilde{x}_\pi^{MP}$	35.622*** (9.892)	31.954*** (8.634)	20.297*** (4.258)	17.350*** (3.626)	11.168*** (2.053)	9.707*** (1.744)
$\tilde{x}_x^{MP}$	-19.889* (10.627)	-16.551* (9.327)	-7.608* (4.250)	-6.497* (3.682)	-2.580** (1.243)	-2.160* (1.121)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the civilian unemployment rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.17: Impact of monetary policy on excess bond returns with alternative VAR specification [2] – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-24.222*** (1.999)	-54.280*** (4.049)	-25.728*** (1.427)	-34.502*** (1.960)	-17.381*** (0.709)	-19.113*** (0.789)
$\tilde{x}_{r^i}^{MP}$	10.554** (4.625)	17.922** (7.532)	11.228** (3.915)	14.764*** (4.968)	6.609* (3.386)	7.176* (3.660)
$\tilde{x}_\pi^{MP}$	17.371 (11.754)	44.258** (18.447)	16.941** (7.015)	23.765*** (8.448)	11.914** (4.563)	13.464*** (4.722)
$\tilde{x}_x^{MP}$	-3.703 (12.333)	-7.900 (21.634)	-2.441 (6.189)	-4.028 (8.376)	-1.141 (2.461)	-1.527 (2.665)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.16 notes.

Table C2.18: Impact of monetary policy on excess bond returns with alternative VAR specification [2] – MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.791*** (0.290)	-0.883 (1.113)	0.778*** (0.120)	0.712 (0.549)	0.329*** (0.041)	0.471** (0.210)
$\tilde{x}_{r^i}^{MP}$	-0.334 (0.228)	1.105 (0.744)	-0.158 (0.152)	0.415 (0.363)	0.112 (0.103)	-0.016 (0.203)
$\tilde{x}_{\pi}^{MP}$	-2.117*** (0.614)	0.865 (2.292)	-1.220*** (0.278)	-0.757 (0.862)	-0.635*** (0.130)	-0.389 (0.367)
$\tilde{x}_x^{MP}$	1.660* (0.850)	-1.087 (1.755)	0.600* (0.318)	-0.370 (0.485)	0.194** (0.074)	-0.067 (0.128)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.16 notes.

Table C2.19: Impact of monetary policy on excess bond returns with alternative VAR specification [2] – Unexpected MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	1.051*** (0.104)	1.203** (0.573)	1.185*** (0.049)	1.089*** (0.331)	0.508*** (0.019)	0.339** (0.144)
$\tilde{x}_{r^i}^{MP}$	0.056 (0.179)	0.093 (0.340)	0.011 (0.177)	-0.359 (0.277)	0.193 (0.154)	-0.421** (0.183)
$\tilde{x}_{\pi}^{MP}$	-2.170*** (0.651)	1.561 (1.217)	-1.659*** (0.403)	0.407 (0.628)	-0.898*** (0.227)	0.438 (0.300)
$\tilde{x}_x^{MP}$	1.062 (0.698)	-2.856** (1.287)	0.463 (0.344)	-1.138** (0.551)	0.197* (0.102)	-0.356** (0.160)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.16 notes.

Table C2.20: Impact of monetary policy on excess bond returns with alternative VAR specification [3] – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-18.963*** (5.835)	-18.966*** (5.237)	-14.380*** (3.133)	-13.010*** (2.901)	-8.125*** (1.367)	-7.705*** (1.312)
$\tilde{x}_{r^i}^{MP}$	0.857 (1.781)	1.306 (1.555)	-1.529 (1.246)	-0.812 (1.031)	-2.150 (1.507)	-1.388 (1.210)
$\tilde{x}_\pi^{MP}$	35.285*** (8.262)	31.220*** (7.311)	22.443*** (3.694)	18.680*** (3.415)	12.770*** (1.815)	10.712*** (1.508)
$\tilde{x}_x^{MP}$	-17.178** (8.642)	-13.560* (7.047)	-6.533* (3.623)	-4.858 (3.078)	-2.494** (1.238)	-1.619** (1.109)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the Chicago Fed National Activity Index and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.21: Impact of monetary policy on excess bond returns with alternative VAR specification [3] – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-23.299*** (2.747)	-51.647*** (5.510)	-24.104*** (1.657)	-32.266*** (2.459)	-16.348*** (0.833)	-17.850*** (0.956)
$\tilde{x}_{r^i}^{MP}$	-2.182** (0.972)	-2.231 (1.534)	-0.760 (1.447)	-0.651 (2.052)	1.811 (2.796)	1.810 (3.323)
$\tilde{x}_\pi^{MP}$	14.389*** (4.890)	44.131*** (6.105)	22.134*** (4.866)	32.953*** (5.810)	15.085*** (4.008)	17.803*** (4.437)
$\tilde{x}_x^{MP}$	11.092** (4.257)	9.746 (8.290)	2.730 (4.265)	-0.035 (5.957)	-0.548 (2.311)	-1.763 (2.604)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.20 notes.

Table C2.22: Impact of monetary policy on excess bond returns with alternative VAR specification [3] – MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.566	-0.644	0.608***	0.968*	0.255***	0.611***
	(0.451)	(1.009)	(0.159)	(0.538)	(0.052)	(0.216)
$\tilde{x}_{r^i}^{MP}$	0.114	0.549**	0.226*	-0.456**	0.261*	-0.456**
	(0.139)	(0.262)	(0.130)	(0.199)	(0.136)	(0.192)
$\tilde{x}_\pi^{MP}$	-2.329***	0.653	-1.599***	-0.296	-0.825***	0.034
	(0.546)	(1.487)	(0.299)	(0.752)	(0.179)	(0.363)
$\tilde{x}_x^{MP}$	1.649*	0.540	0.765**	-0.215	0.309***	-0.189
	(0.869)	(0.910)	(0.339)	(0.382)	(0.101)	(0.130)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.20 notes.

Table C2.23: Impact of monetary policy on excess bond returns with alternative VAR specification [3] – Unexpected MB change

	10-year bonds		**5-year bonds**		**2-year bonds**	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.978***	1.477**	1.119***	1.274***	0.479***	0.413***
	(0.143)	(0.687)	(0.063)	(0.356)	(0.023)	(0.149)
$\tilde{x}_{r^i}^{MP}$	0.218	-0.044	0.301	-0.314	0.285	-0.535**
	(0.182)	(0.237)	(0.200)	(0.220)	(0.222)	(0.203)
$\tilde{x}_\pi^{MP}$	-2.544**	1.272	-2.092***	0.170	-1.091***	0.449
	(0.940)	(1.211)	(0.573)	(0.776)	(0.332)	(0.358)
$\tilde{x}_x^{MP}$	1.348	-2.705**	0.673	-1.129**	0.326**	-0.328*
	(0.897)	(1.108)	(0.428)	(0.545)	(0.144)	(0.169)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.20 notes.

Table C2.24: Impact of monetary policy on excess bond returns with alternative VAR specification [4] – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-21.102*** (4.647)	-22.970*** (4.684)	-15.615*** (2.962)	-14.930*** (2.832)	-8.357*** (1.356)	-8.111*** (1.291)
$\tilde{x}_{r^i}^{MP}$	2.193 (1.931)	4.003* (2.350)	-0.468 (1.282)	1.582 (1.829)	-1.270 (1.232)	0.730 (1.335)
$\tilde{x}_\pi^{MP}$	29.886*** (7.858)	21.440** (9.521)	19.765*** (3.788)	13.267*** (5.053)	11.455*** (1.524)	7.778*** (1.853)
$\tilde{x}_x^{MP}$	-10.977* (6.620)	-2.472 (8.519)	-3.682 (3.204)	0.082 (4.018)	-1.828 (1.166)	-0.398 (1.233)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, the Chicago Fed Adjusted National Financial Conditions Index and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.25: Impact of monetary policy on excess bond returns with alternative VAR specification [4] – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-27.841*** (3.035)	-54.337*** (3.379)	-27.168*** (1.769)	-34.897*** (1.776)	-18.018*** (0.889)	-19.337*** (0.735)
$\tilde{x}_{r^i}^{MP}$	-2.078* (1.149)	-0.891 (1.878)	-0.737 (2.154)	0.964 (2.066)	2.216 (3.735)	3.822 (3.288)
$\tilde{x}_\pi^{MP}$	20.157*** (4.616)	36.534*** (8.092)	25.690*** (6.328)	28.266*** (6.107)	16.839*** (5.352)	15.258*** (4.436)
$\tilde{x}_x^{MP}$	9.762** (4.597)	18.694** (7.541)	2.215 (5.006)	5.667 (5.649)	-1.037 (2.638)	0.258 (2.624)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.24 notes.

Table C2.26: Impact of monetary policy on excess bond returns with alternative VAR specification [4] – MB change

ΔMB	10-year bonds 1985:1 – 2014:2	10-year bonds 1985:1 – 2007:7	5-year bonds 1985:1 – 2014:2	5-year bonds 1985:1 – 2007:7	2-year bonds 1985:1 – 2014:2	2-year bonds 1985:1 – 2007:7
$\tilde{x}_n^{MP}$	0.475	-0.720	0.661***	0.849	0.287***	0.537**
	(0.426)	(1.008)	(0.153)	(0.521)	(0.053)	(0.207)
$\tilde{x}_{r^i}^{MP}$	0.229*	-0.178	0.323***	-0.080	0.340***	-0.098
	(0.121)	(0.314)	(0.114)	(0.253)	(0.118)	(0.241)
$\tilde{x}_\pi^{MP}$	-2.298***	-1.368	-1.598***	-1.526**	-0.851***	-0.641*
	(0.390)	(1.272)	(0.230)	(0.718)	(0.136)	(0.353)
$\tilde{x}_x^{MP}$	1.595**	2.266*	0.615**	0.757	0.225***	-0.203
	(0.703)	(1.233)	(0.280)	(0.533)	(0.074)	(0.158)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.24 notes.

Table C2.27: Impact of monetary policy on excess bond returns with alternative VAR specification [4] – Unexpected MB change

ΔMB^U	10-year bonds 1985:1 – 2014:2	10-year bonds 1985:1 – 2007:7	5-year bonds 1985:1 – 2014:2	5-year bonds 1985:1 – 2007:7	2-year bonds 1985:1 – 2014:2	2-year bonds 1985:1 – 2007:7
$\tilde{x}_n^{MP}$	0.723**	0.440	1.037***	0.702*	0.455***	0.206
	(0.304)	(0.814)	(0.127)	(0.427)	(0.044)	(0.169)
$\tilde{x}_{r^i}^{MP}$	0.608*	0.325	0.657*	-0.004	0.604*	-0.265
	(0.339)	(0.327)	(0.353)	(0.347)	(0.355)	(0.296)
$\tilde{x}_\pi^{MP}$	-3.600**	0.469	-2.739***	-0.270	-1.490***	0.179
	(1.478)	(1.226)	(0.891)	(0.921)	(0.477)	(0.437)
$\tilde{x}_x^{MP}$	2.269	-1.234	1.045	-0.428	0.431**	-0.120
	(1.458)	(1.267)	(0.690)	(0.723)	(0.194)	(0.209)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.24 notes.

Table C2.28: Impact of monetary policy on excess bond returns – Romer and Romer policy shock

	10-year bonds	5-year bonds	2-year bonds
ΔRR	**1985:1 – 2007:7**	**1985:1 – 2007:7**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-27.695***	-17.632***	-8.729***
	(4.765)	(3.108)	(1.487)
$\tilde{x}_{r^i}^{MP}$	2.208	-0.113	-0.779
	(1.641)	(1.006)	(1.181)
$\tilde{x}_\pi^{MP}$	39.123***	22.428***	11.240***
	(7.343)	(3.226)	(1.421)
$\tilde{x}_x^{MP}$	-13.635**	-4.683	-1.733*
	(6.788)	(3.036)	(1.027)

Notes: This table reports the impact of a monetary policy shock as measured by Romer and Romer (2004) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate, and the relative bill rate. The pre-crisis period (1985:1 – 2007:7) results are reported for each bond maturity. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.29: Variance decomposition for excess bond returns – VAR(3)

	10-year bonds		**5-year bonds**		**2-year bonds**	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var(\tilde{x}_\pi)$	0.980***	1.042***	1.532**	1.386**	2.339*	1.847**
	(0.362)	(0.370)	(0.675)	(0.576)	(1.216)	(0.892)
$2Cov(\tilde{x}_\pi, \tilde{x}_{r^i})$	-0.071	-0.049	-0.226	-0.120	-1.078	-0.429
	(0.094)	(0.069)	(0.272)	(0.155)	(1.120)	(0.658)
$2Cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.058	-0.122	-0.582	-0.493	-1.361	-0.997
	(0.387)	(0.429)	(0.684)	(0.656)	(0.838)	(0.754)
$Var(\tilde{x}_{r^i})$	0.016*	0.008	0.041	0.015	0.371	0.101
	(0.009)	(0.005)	(0.030)	(0.012)	(0.242)	(0.097)
$2Cov(\tilde{x}_{r^i}, \tilde{x}_x)$	-0.020	-0.039	0.076	0.016	0.463	0.198
	(0.052)	(0.040)	(0.107)	(0.073)	(0.360)	(0.239)
$Var(\tilde{x}_x)$	0.153	0.161	0.159	0.195	0.266	0.279
	(0.126)	(0.173)	(0.165)	(0.207)	(0.179)	(0.203)
$R^2(\tilde{x}_\pi)$	0.856***	0.877***	0.831***	0.841***	0.536**	0.696***
	(0.114)	(0.137)	(0.131)	(0.148)	(0.194)	(0.179)
$R^2(\tilde{x}_{r^i})$	0.055	0.175	0.029	0.086	0.011	0.002
	(0.096)	(0.184)	(0.113)	(0.171)	(0.091)	(0.088)
$R^2(\tilde{x}_x)$	0.084	0.040	0.056	0.010	0.126	0.051
	(0.222)	(0.160)	(0.244)	(0.096)	(0.252)	(0.159)

Notes: This table reports the variance decomposition of the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_x$) and the covariances between these three components. News components are extracted from a VAR(3) model where the state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample (1985:1 – 2014:2) and pre-crisis period (1985:1 – 2007:7) results, respectively. R^2 values are obtained from regressions of unexpected excess returns on each news component. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.30: Impact of monetary policy on excess bond returns with VAR(3) – FFR change

ΔFFR	10-year bonds 1985:1 – 2014:2	10-year bonds 1985:1 – 2007:7	5-year bonds 1985:1 – 2014:2	5-year bonds 1985:1 – 2007:7	2-year bonds 1985:1 – 2014:2	2-year bonds 1985:1 – 2007:7
$\tilde{x}_n^{MP}$	-22.902***	-22.432***	-14.481***	-12.573***	-7.517***	-6.898***
	(5.215)	(5.290)	(3.177)	(3.089)	(1.400)	(1.362)
$\tilde{x}_{r^i}^{MP}$	0.618	1.377	-2.345	-1.263	-2.570	-1.440
	(1.789)	(1.543)	(1.521)	(1.169)	(1.980)	(1.617)
$\tilde{x}_\pi^{MP}$	34.913***	31.089***	25.220***	20.762***	13.963***	11.428***
	(8.991)	(7.698)	(5.274)	(4.330)	(2.485)	(1.984)
$\tilde{x}_x^{MP}$	-12.630	-10.034	-8.395*	-6.926*	-3.876**	-3.090**
	(8.756)	(7.624)	(4.882)	(4.073)	(1.405)	(1.217)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(3) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.31: Impact of monetary policy on excess bond returns with VAR(3) – Unexpected FFR change

ΔFFR^U	10-year bonds 1989:2 – 2014:2	10-year bonds 1989:2 – 2007:7	5-year bonds 1989:2 – 2014:2	5-year bonds 1989:2 – 2007:7	2-year bonds 1989:2 – 2014:2	2-year bonds 1989:2 – 2007:7
$\tilde{x}_n^{MP}$	-30.890***	-59.543***	-28.425***	-36.490***	-18.091***	-19.443***
	(4.348)	(4.924)	(2.335)	(2.495)	(0.966)	(0.977)
$\tilde{x}_{r^i}^{MP}$	-1.745	-1.534	-0.588	-0.748	1.549	1.550
	(1.081)	(1.534)	(2.010)	(2.398)	(4.663)	(4.748)
$\tilde{x}_\pi^{MP}$	25.186***	53.557***	31.207***	40.711***	20.411***	21.811***
	(5.876)	(7.563)	(7.460)	(8.132)	(6.877)	(6.728)
$\tilde{x}_x^{MP}$	7.449	7.520	-2.194	-3.473	-3.869	-3.918
	(4.836)	(8.295)	(6.075)	(7.483)	(3.503)	(3.550)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.30 notes.

Table C2.32: Impact of monetary policy on excess bond returns with VAR(3) – MB change

ΔMB	10-year bonds 1985:1 – 2014:2	10-year bonds 1985:1 – 2007:7	5-year bonds 1985:1 – 2014:2	5-year bonds 1985:1 – 2007:7	2-year bonds 1985:1 – 2014:2	2-year bonds 1985:1 – 2007:7
$\tilde{x}_n^{MP}$	0.612*	-0.009	0.673***	0.950	0.282***	0.527**
	(0.354)	(1.219)	(0.150)	(0.625)	(0.055)	(0.243)
$\tilde{x}_{r^i}^{MP}$	0.207*	-0.273	0.343**	-0.341*	0.349***	-0.418**
	(0.124)	(0.215)	(0.135)	(0.193)	(0.143)	(0.191)
$\tilde{x}_\pi^{MP}$	-1.927***	0.781	-1.607***	-0.389	-0.887***	-0.066
	(0.673)	(1.745)	(0.434)	(0.936)	(0.211)	(0.423)
$\tilde{x}_x^{MP}$	1.108*	-0.499	0.592	-0.220	0.256***	-0.043
	(0.644)	(0.747)	(0.359)	(0.421)	(0.098)	(0.173)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.30 notes.

Table C2.33: Impact of monetary policy on excess bond returns with VAR(3) – Unexpected MB change

ΔMB^U	10-year bonds 1985:1 – 2014:2	10-year bonds 1985:1 – 2007:7	5-year bonds 1985:1 – 2014:2	5-year bonds 1985:1 – 2007:7	2-year bonds 1985:1 – 2014:2	2-year bonds 1985:1 – 2007:7
$\tilde{x}_n^{MP}$	0.719**	1.668**	1.056***	1.308***	0.458***	0.415***
	(0.329)	(0.699)	(0.136)	(0.373)	(0.045)	(0.154)
$\tilde{x}_{r^i}^{MP}$	0.317**	0.021	0.408**	-0.434	0.367*	-0.745***
	(0.144)	(0.265)	(0.186)	(0.271)	(0.217)	(0.264)
$\tilde{x}_\pi^{MP}$	-1.921**	1.061	-2.032***	0.410	-1.110***	0.837*
	(0.873)	(1.493)	(0.623)	(1.069)	(0.328)	(0.500)
$\tilde{x}_x^{MP}$	0.885	-2.750**	0.568	-1.284	0.285**	-0.507*
	(0.674)	(1.290)	(0.447)	(0.814)	(0.139)	(0.258)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.30 notes.

Table C2.34: Variance decomposition for excess bond returns – VAR(6)

	10-year bonds		5-year bonds		2-year bonds	
	1985:1–2014:2	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**	**1985:1–2014:2**	**1985:1–2007:7**
$Var(\tilde{x}_{\pi})$	1.329** (0.588)	1.744* (1.040)	1.836** (0.898)	1.887* (1.116)	2.877** (1.439)	2.135* (1.190)
$2Cov(\tilde{x}_{\pi}, \tilde{x}_{r^i})$	-0.109 (0.172)	-0.143 (0.211)	-0.410 (0.436)	-0.297 (0.435)	-1.909 (1.490)	-0.933 (1.168)
$2Cov(\tilde{x}_{\pi}, \tilde{x}_{x})$	-0.458 (0.664)	-1.069 (1.383)	-0.958 (0.945)	-1.301 (1.389)	-1.723* (0.945)	-1.371 (0.910)
$Var(\tilde{x}_{r^i})$	0.031* (0.018)	0.014 (0.012)	0.099 (0.067)	0.040 (0.043)	0.719* (0.419)	0.318 (0.279)
$2Cov(\tilde{x}_{r^i}, \tilde{x}_{x})$	0.018 (0.096)	0.058 (0.123)	0.199 (0.203)	0.203 (0.244)	0.724 (0.475)	0.507 (0.441)
$Var(\tilde{x}_{x})$	0.189 (0.184)	0.396 (0.454)	0.234 (0.240)	0.468 (0.450)	0.311* (0.187)	0.344 (0.212)
$R^2(\tilde{x}_{\pi})$	0.822*** (0.129)	0.743*** (0.191)	0.723*** (0.180)	0.627*** (0.213)	0.391* (0.202)	0.453** (0.208)
$R^2(\tilde{x}_{r^i})$	0.006 (0.044)	0.059 (0.154)	0.000 (0.015)	0.001 (0.027)	0.022 (0.110)	0.034 (0.145)
$R^2(\tilde{x}_{x})$	0.005 (0.070)	0.030 (0.140)	0.090 (0.288)	0.014 (0.090)	0.114 (0.219)	0.022 (0.092)

Notes: This table reports the variance decomposition of the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds into the variances of inflation news ($\tilde{x}_{\pi}$), real interest rate news ($\tilde{x}_{r^i}$), risk premium news ($\tilde{x}_{x}$) and the covariances between these three components. News components are extracted from a VAR(6) model where the state vector contains the first difference in 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. See also Table C2.29 notes.

Table C2.35: Impact of monetary policy on excess bond returns with VAR(6) – FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	-22.725*** (5.538)	-20.819*** (5.681)	-13.819*** (3.285)	-12.078*** (3.238)	-7.422*** (1.395)	-6.763*** (1.382)
$\tilde{x}_{r^i}^{MP}$	-0.659 (2.593)	0.771 (2.135)	-3.990 (2.456)	-2.142 (1.823)	-3.428 (2.161)	-1.620 (1.732)
$\tilde{x}_\pi^{MP}$	39.088*** (11.067)	34.119*** (9.223)	28.302*** (6.455)	22.246*** (4.901)	15.072*** (2.591)	11.446*** (2.012)
$\tilde{x}_x^{MP}$	-17.703* (10.616)	-14.071 (8.998)	-10.493* (5.643)	-8.025* (4.425)	-4.222*** (1.356)	-3.063*** (1.123)

Notes: This table reports the impact of a change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(6) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.36: Impact of monetary policy on excess bond returns with VAR(6) – Unexpected FFR change

	10-year bonds		5-year bonds		2-year bonds	
ΔFFR^U	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**	**1989:2 – 2014:2**	**1989:2 – 2007:7**
$\tilde{x}_n^{MP}$	-32.798*** (7.133)	-59.358*** (7.060)	-28.720*** (3.665)	-36.332*** (3.649)	-17.600*** (1.437)	-18.986*** (1.467)
$\tilde{x}_{r^i}^{MP}$	-0.853 (1.980)	-1.197 (2.781)	0.205 (3.391)	-0.740 (3.565)	2.515 (5.826)	1.850 (5.546)
$\tilde{x}_\pi^{MP}$	32.868*** (10.459)	63.819*** (11.812)	32.903*** (10.594)	43.541*** (10.172)	18.943*** (8.268)	21.400*** (7.432)
$\tilde{x}_x^{MP}$	0.783 (7.138)	-3.264 (11.900)	-4.388 (7.562)	-6.469 (8.831)	-3.857 (3.599)	-4.264 (3.541)

Notes: This table reports the impact of an unexpected change in the federal funds rate (FFR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). Due to data availability on FFR futures, the full sample that is used for the estimations of monetary policy effects commences on 1989:2. See also Table C2.35 notes.

Table C2.37: Impact of monetary policy on excess bond returns with VAR(6) – MB change

ΔMB	10-year bonds		5-year bonds		2-year bonds	
	1985:1 – 2014:2	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.583 (0.470)	0.097 (1.236)	0.694*** (0.197)	0.998 (0.650)	0.301*** (0.067)	0.561** (0.257)
$\tilde{x}_{r^i}^{MP}$	0.404* (0.239)	-0.472 (0.339)	0.590** (0.252)	-0.706** (0.351)	0.560** (0.203)	-0.793** (0.327)
$\tilde{x}_\pi^{MP}$	-2.509*** (1.196)	1.584 (2.120)	-2.199*** (0.723)	0.402 (1.210)	-1.232*** (0.299)	0.472 (0.553)
$\tilde{x}_x^{MP}$	1.522* (0.918)	-1.208 (1.078)	0.915* (0.524)	-0.695 (0.655)	0.371*** (0.126)	-0.239 (0.221)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.35 notes.

Table C2.38: Impact of monetary policy on excess bond returns with VAR(6) – Unexpected MB change

ΔMB^U	10-year bonds		5-year bonds		2-year bonds	
	1985:1 – 2014:2	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.665 (0.406)	1.523* (0.873)	1.067*** (0.166)	1.251*** (0.454)	0.472*** (0.054)	0.386** (0.181)
$\tilde{x}_{r^i}^{MP}$	0.514** (0.253)	-0.304 (0.469)	0.660** (0.294)	-0.944* (0.527)	0.583** (0.271)	-1.165*** (0.410)
$\tilde{x}_\pi^{MP}$	-2.539** (1.296)	2.133 (2.086)	-2.675*** (0.857)	1.513 (1.506)	-1.468*** (0.391)	1.485** (0.645)
$\tilde{x}_x^{MP}$	1.360 (0.965)	-3.351* (1.796)	0.948 (0.609)	-1.820 (1.125)	0.413** (0.159)	-0.707** (0.296)

Notes: This table reports the impact of a change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). See also Table C2.35 notes.

Table C2.39: Impact of monetary policy on excess bond returns – Unexpected MB change – alternative measure [1]

	10-year bonds		5-year bonds		2-year bonds	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.393*** (0.124)	-2.531*** (0.383)	0.828*** (0.059)	-0.855*** (0.221)	0.369*** (0.024)	-0.325*** (0.090)
$\tilde{x}_{r^i}^{MP}$	0.355*** (0.104)	0.156 (0.269)	0.360*** (0.120)	-0.134 (0.206)	0.304** (0.142)	-0.356** (0.178)
$\tilde{x}_\pi^{MP}$	-1.265** (0.502)	5.125*** (1.252)	-1.373*** (0.351)	2.314*** (0.611)	-0.799*** (0.205)	1.166*** (0.279)
$\tilde{x}_x^{MP}$	0.517 (0.447)	-2.749** (1.194)	0.185 (0.252)	-1.324** (0.564)	0.127 (0.092)	-0.485** (0.186)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The model used to extract unexpected changes in MB includes seven lags of its own, seven lags of unemployment measure and seven lags of the Chicago Fed Adjusted National Financial Conditions Index. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.40: Impact of monetary policy on excess bond returns – Unexpected MB change – alternative measure [2]

	10-year bonds		5-year bonds		2-year bonds	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	1.096*** (0.197)	2.146*** (0.515)	1.270*** (0.084)	1.322*** (0.302)	0.556*** (0.023)	0.461*** (0.137)
$\tilde{x}_{r^i}^{MP}$	0.194** (0.093)	-0.095 (0.236)	0.219** (0.110)	-0.387** (0.182)	0.179 (0.129)	-0.555*** (0.161)
$\tilde{x}_\pi^{MP}$	-2.041*** (0.430)	-0.010 (1.332)	-1.774*** (0.337)	-0.240 (0.672)	-0.887*** (0.191)	0.310 (0.289)
$\tilde{x}_x^{MP}$	0.751* (0.433)	-2.040** (0.859)	0.286 (0.245)	-0.695* (0.372)	0.152* (0.085)	-0.217* (0.116)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The model used to extract unexpected changes in MB includes seven lags of its own and seven lags of the first difference in log industrial production index. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.41: Impact of monetary policy on excess bond returns – Unexpected MB change – alternative measure [3]

	10-year bonds		5-year bonds		2-year bonds	
ΔMB^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	1.193*** (0.169)	2.201*** (0.384)	1.325*** (0.067)	1.190*** (0.204)	0.590*** (0.018)	0.456*** (0.097)
$\tilde{x}_{r^i}^{MP}$	0.205** (0.085)	0.094 (0.223)	0.214* (0.109)	-0.204 (0.164)	0.158 (0.135)	-0.396*** (0.141)
$\tilde{x}_\pi^{MP}$	-1.940*** (0.384)	-0.033 (1.205)	-1.754*** (0.326)	-0.278 (0.599)	-0.884*** (0.193)	0.119 (0.258)
$\tilde{x}_x^{MP}$	0.542 (0.398)	-2.262*** (0.824)	0.215 (0.255)	-0.709* (0.366)	0.137 (0.092)	-0.179 (0.111)

Notes: This table reports the impact of an unexpected change in log monetary base (MB) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). The model used to extract unexpected changes in MB includes nine lags of its own, nine lags of the first difference in log industrial production index and nine lags of the first difference in 3-month Treasury bill rate. See also Table C2.39 notes.

Table C2.42: Impact of monetary policy on excess bond returns – TR change

	10-year bonds		5-year bonds		2-year bonds	
ΔTR	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.166** (0.077)	0.307** (0.142)	0.300*** (0.029)	0.471*** (0.067)	0.149*** (0.011)	0.216*** (0.025)
$\tilde{x}_{r^i}^{MP}$	0.058** (0.026)	-0.071 (0.058)	0.083*** (0.027)	-0.108** (0.051)	0.078** (0.034)	-0.158*** (0.052)
$\tilde{x}_\pi^{MP}$	-0.551*** (0.101)	0.209 (0.344)	-0.488*** (0.062)	-0.133 (0.173)	-0.277*** (0.043)	0.020 (0.085)
$\tilde{x}_x^{MP}$	0.327** (0.151)	-0.446** (0.214)	0.106 (0.069)	-0.230* (0.119)	0.050** (0.024)	-0.078* (0.046)

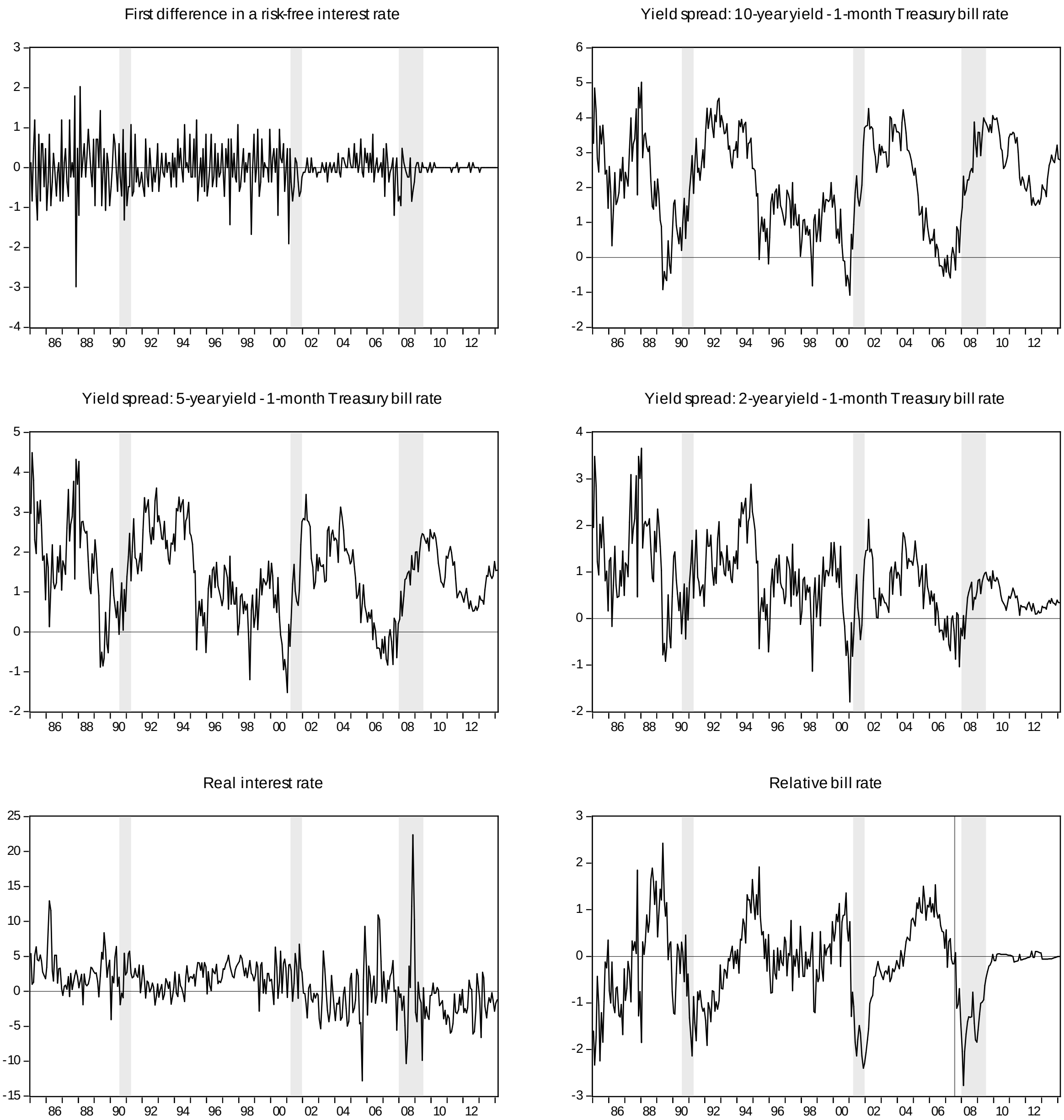
Notes: This table reports the impact of a change in log adjusted St. Louis total reserves (TR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). News components are extracted from a VAR(1) model estimated over the full sample period (1985:1 – 2014:2). The state vector contains the first difference of the 1-month Treasury bill rate, the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill, the real interest rate and the relative bill rate. The first and second column for each bond maturity report the full sample and pre-crisis period (1985:1 – 2007:7) results, respectively. The standard errors reported in parentheses are computed using the delta method. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table C2.43: Impact of monetary policy on excess bond returns – Unexpected TR change

	10-year bonds		5-year bonds		2-year bonds	
ΔTR^U	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**	**1985:1 – 2014:2**	**1985:1 – 2007:7**
$\tilde{x}_n^{MP}$	0.093** (0.041)	0.483*** (0.103)	0.334*** (0.019)	0.343*** (0.059)	0.180*** (0.007)	0.123*** (0.024)
$\tilde{x}_{r^i}^{MP}$	0.042*** (0.010)	-0.015 (0.06)	0.036 (0.023)	-0.086* (0.051)	0.006 (0.037)	-0.139*** (0.047)
$\tilde{x}_\pi^{MP}$	-0.130*** (0.047)	0.198 (0.315)	-0.339*** (0.076)	0.003 (0.173)	-0.193*** (0.053)	0.096 (0.079)
$\tilde{x}_x^{MP}$	-0.005 (0.043)	-0.665*** (0.251)	-0.031 (0.058)	-0.261** (0.121)	0.007 (0.027)	-0.080** (0.038)

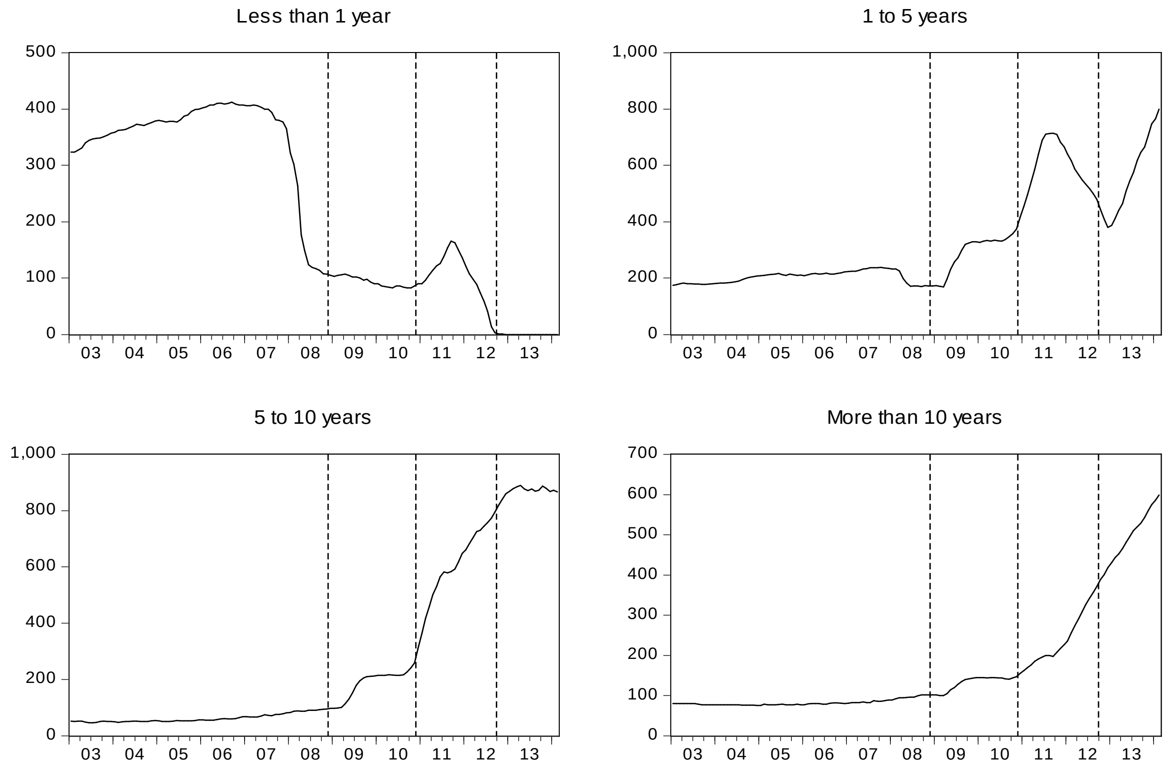
Notes: This table reports the impact of an unexpected change in log adjusted St. Louis total reserves (TR) on the unexpected excess returns of 10-, 5-, and 2-year Treasury bonds ($\tilde{x}_n$), inflation news ($\tilde{x}_\pi$), real interest rate news ($\tilde{x}_{r^i}$) and risk premium news ($\tilde{x}_x$). The model used to extract unexpected changes in TR includes seven lags of its own and seven lags of the unemployment measure as defined in Section 4.3. See also Table C2.41 notes.

Figure C2.1: VAR state variables



Notes: This figure plots the variables used for the benchmark VAR estimations over the full sample period 1985:1 – 2014:2; the first difference of the 1-month Treasury bill rate (Δy_1), the yield spread between 10-, 5- and 2-year Treasury bonds and the 1-month Treasury bill (s_n), the real interest rate (r^i) and the relative bill rate (*rb*). All variables are expressed in percentages per annum on continuously compounded basis. Shaded areas denote US recessions as classified by NBER business cycle dates.

Figure C2.2: US Treasury securities held by the Fed



Notes: This figure plots the Federal Reserve's US Treasury securities holdings (in $bn). The upper left panel plots the holdings of securities with maturity of less than one year; the upper right panel plots the holdings of securities with maturity between one and five years; the lower left panel plots the holdings of securities with maturity between five and ten years; the lower right panel plots the holdings of securities with maturity of more than ten years. The three dashed vertical lines denote the announcements of the first round of quantitative easing (QE1, 2008:11), the second round (QE2, 2010:11) and the third round (QE3, 2012:9). Data is obtained from FRED database.

Chapter 3: US monetary policy and stock prices: revisiting the size and value effects

3.1 Introduction

Finance practitioners and academics largely agree that monetary policy has key implications for financial markets and that there may be a role for asset prices in the monetary policy reaction function. In response to the global financial crisis, the Federal Reserve (Fed) has expanded its toolkit with unconventional monetary policies, such as liquidity facilities and large-scale asset purchases known as quantitative easing. As economy improved considerably, the purchases of longer-term assets were discontinued in October 2014. The federal funds rate (FFR) target was finally raised for the first time in almost a decade in December 2015 bringing conventional monetary policy back to the spotlight.

The empirical literature investigating monetary policy effects on stock prices in the United States goes back to the 1970s. In these early studies it is already acknowledged that the causal relationship between stock prices and monetary policy may run in both directions (Keran, 1971, Cooper, 1974). Since the global financial crisis, the old academic debate on the appropriate response of monetary policy to financial developments has been revived (Bernanke and Gertler, 1999; Cecchetti et al. 2000; Kuttner, 2011). Given that stock prices play an important role in the monetary policy transmission, it is vital to gain a thorough understanding of how monetary policy interacts with the stock market (Mishkin, 2001; Bjornland and Jacobsen, 2013). Consequently, recent empirical literature is increasingly more focused on the interdependence between monetary policy behaviour and developments in stock prices or, more generally, financial markets.

Monetary policy may affect stock prices through its impact on expected future net cash flows and the discount rate, i.e. the sum of a risk-free interest rate and a risk premium (Homa and Jaffee, 1971; Smirlock and Yawitz, 1985). With respect to empirical evidence, contractionary monetary policy is typically associated with a significant decline in the stock market. This effect tends to be more pronounced in bad economic times and bear markets (Basistha and Kurov, 2008; Jansen and Tsai, 2010, Kurov, 2010). The transmission mechanism of monetary policy implies that some firms may be more exposed to changing monetary conditions than others. It is generally found that monetary policy actions have a stronger impact on small and value stocks as compared to large and growth stocks, i.e. the size and value effects of monetary policy (Thorbecke, 1997; Ehrmann and Fratzscher, 2004; Guo, 2004; Basistha and Kurov, 2008; Jansen and Tsai, 2010; Tsai, 2011; Kontonikas and Kostakis, 2013). As Lioui and Maio (2014) demonstrate, value stocks earn an additional risk premium relative to growth stocks with respect to a monetary policy risk factor. The strand of literature examining monetary policy effects on stocks relates to the stock market anomalies analysed in the cross-sectional asset pricing literature (Fama and French, 1993; 1995) as the monetary policy factor may help to explain stock return differentials across firms with different characteristics.

Nevertheless, the evidence of the differential impact of monetary policy on small and value stocks versus large and growth stocks in the US appears to be weaker and mixed since the 1980s (Guo, 2004; Tsai, 2011; Kontonikas and Kostakis, 2013; Maio, 2014). Furthermore, the majority of empirical work is typically focused only on one side of the potentially bi-directional relationship between monetary policy and stock prices. In two separate studies, Rigobon and Sack (2003; 2004) show that the bi-directional causality exists between US monetary policy and the stock market. Consequently, more recent research examines this simultaneous relationship (Bjornland and Leitemo, 2009; D'Amico and Farka, 2011; Bouakez, Essid and Normandin, 2013). On the other hand, these empirical studies that take into account the simultaneity between monetary policy and stock prices are typically focused on the stock market level. This type of analysis at the disaggregated level is very scant, especially using lower frequency data.

To fill the gap, this chapter investigates the interaction between conventional US monetary policy and real stock prices at both the aggregate market and portfolio levels in the spirit of Bjornland and Leitemo (2009). Following the first steps towards monetary policy normalisation in the US, this study is useful for future policy decision making with respect to standard monetary policy tools. The structural vector autoregression (SVAR) model is identified using a combination of standard short-run zero restrictions and one long-run restriction that implies monetary policy neutrality. It is assumed that monetary policy has no effect on real stock prices in the long run leaving the short-run relationship between real stock returns and the policy rate unconstrained.

Several important contributions to the existing literature are made. Firstly and most importantly, this chapter revisits the role of US monetary policy in explaining the size and value stock market anomalies using a single model that allows for a fully simultaneous interaction between the policy interest rate and real stock returns. Equivalently, the empirical analysis provides the insight into the policy reaction to stock price developments at the market level and stock portfolio level, while taking into account a contemporaneous stock price response to policy shocks. Secondly, the model specification as in Bjornland and Leitemo (2009) is augmented in line with the recommendations by Brissimis and Magginas (2006). Two forward-looking variables, market-based expectations about the level of monetary policy rate and a composite leading indicator of economic activity, are included into the otherwise standard SVAR model for the analysis of monetary policy effects. This considerably improves the specification of the monetary policy reaction function and generates a sharper measure of policy shocks. Finally, the main empirical analysis is conducted over the sample period 1994:2 – 2007:7 that is not a standard choice in the SVAR literature. The motivation for the start of the sample stems from the significant changes in the Fed's communication of policy decisions implemented at that time. More transparent and more predictable policy conduct may have had an impact on a stock price response to monetary policy actions. Also, the robustness analysis extends the sample to include the crisis period.

The key findings can be summarised as follows. Firstly, a contractionary monetary policy shock has a strong, negative and statistically significant impact on real stock prices at the market level. Secondly, the results indicate that both the size and value effects of monetary policy prevail to some extent in the period since 1994. Interestingly, the size effect only becomes evident in the second period following the shock. Initially, large stocks respond more negatively to an adverse monetary shock; however, the second-period impulse responses indicate a pronounced decline in small stocks, while large stocks recover to a great extent. The delayed response of smaller stocks to monetary policy shocks could possibly be linked to their relative illiquidity and less frequent trading or the liquidity pull-back and portfolio rebalancing effects. In addition, the learning process of investors may play a role. With respect to the value effect, it appears to be more evident in the case of double-sorted portfolios, i.e. when the size of firms is controlled for. Within each size quintile, the most value stocks are more sensitive to changes in monetary policy conditions than the most growth stocks. Overall, the empirical findings provide some evidence, albeit not very strong, in favour of the credit channel of monetary policy transmission. Finally, the results overall are not supportive of the strong Fed's reaction to stock price developments. After taking into account expectations about future economic activity and the expected level of the policy rate, stock prices seem to provide little relevant information for monetary policymakers beyond what is already included in their information set.

The chapter has the following structure. Sections 3.2 – 3.5 provide the review of related literature. The methodology is explained in Section 3.6. Data description is provided in Section 3.7. Section 3.8 presents some initial findings, while Section 3.9 discusses the main empirical results. The robustness analysis is presented in Section 3.10. Finally, Section 3.11 concludes.

3.2 Monetary policy transmission to stock prices

The price that an investor is willing to pay for a share is equal to the present value of expected future net cash flows (dividends and earnings) discounted at the rate equal to the sum of a risk-free interest rate and a risk premium. It follows that monetary policy may influence current stock prices through its impact on any of these three terms (Homa and Jaffee, 1971; Smirlock and Yawitz, 1985). For instance, monetary policy tightening increases the risk-free rate and, in turn, has a dampening effect on real economy. The expectations of economic slowdown lead to a decline in the expected future earnings of firms and, potentially, to increased uncertainty about future economic and monetary conditions. Consequently, the risk premium required by investors may also rise. Overall, lower expected future cash flows and a higher discount rate imply declining stock prices, falling investment spending, and, eventually, economic contraction (Mishkin, 2001). Given the role that stock prices play in the transmission of monetary policy, it is important to understand how monetary policy decisions influence equity pricing.

The literature distinguishes several channels through which monetary policy may affect stocks. According to the traditional *interest rate* channel, an increase in the policy rate raises the cost of borrowing and dampens the demand for loans. Consequently, a decrease in consumption and investment spending lowers expectations about future net cash flows, hence, current equity prices decline (Bernanke and Gertler, 1995; Ehrmann and Fratzscher, 2004). The *credit* channel works through changes in the supply of funds available to firms. Given imperfect information and credit market frictions, the direct monetary policy effects on market interest rates are amplified through changes in the external finance premium, i.e. the difference between the cost of external funds and internal funds (Gertler and Gilchrist, 1994; Bernanke and Gertler, 1995; Kiyotaki and Moore, 1997). Restricted access to financing reduces investment spending and future cash flows leading to lower equity prices. In addition to the interest rate and credit channels, one may also consider the *risk premium* channel of monetary policy transmission. Bernanke and Kuttner (2005) argue that monetary policy may be associated with changes in the equity (risk) premium. They find that an unexpected increase in the policy rate induces a decline in stock prices due to higher expected future excess returns. Monetary policy tightening may raise the expected risk premium by either increasing the riskiness of firms or by increasing risk aversion among investors due to the higher interest burden and weaker balance sheets of firms.[121]

With respect to the credit channel, there are two mechanisms how monetary policy may affect the external finance premium. The *bank-lending* channel refers to changes in the overall supply of intermediate credit. In restrictive monetary policy conditions, banks reduce their lending and charge higher rates of interest. Hence, funds raised externally become more expensive driving up the external finance premium. Consequently, bank-dependent firms experience the lack of funds and reduce their investment spending that leads to lower expected future net cash flows (Kashyap, Stein and Wilcox, 1993). Alternatively, the *balance-sheet* channel operates through changes in firm's creditworthiness. Following monetary policy tightening, the balance sheet position of a firm deteriorates due to increased interest payments, lower collateral values and lower net worth, and the external finance premium rises. Firstly, monetary policy directly influences debt-servicing expenses and the value of the assets that serve as collateral to obtain external financing (Bernanke and Gertler, 1995). Secondly, it may also affect firm's net worth indirectly through its impact on overall economic activity and sales revenue, i.e. expected future net cash flows.

The above discussion suggests that monetary policy may have a heterogeneous effect on stock prices. For instance, firms operating in cyclical and capital-intensive industries are likely to be more affected by interest rate shocks (Erhmann and Fratzscher, 2004). Similarly, monetary policy actions are expected to have a stronger impact on financially constrained firms. Therefore, small firms, which typically are more bank-dependent and face higher external finance premium than large firms, should also be relatively more exposed to monetary policy risk (Gertler and Gilchrist, 1994; Perez-Quiros and Timmermann, 2000). As alternative measures of the degree of financial distress, various financial ratios may be used. For instance, firms with high earnings-to-price (E/P), book-to-market value (BE/ME), cash-flow-to-price (C/P), and dividend-to-price (D/P) ratios have fewer growth opportunities and are more heavily reliant on cash flows. Such value firms typically are less profitable, experience poor earnings, are more likely to be financially distressed and their stocks tend to be undervalued by the market (Fama and

[121] Alternatively to the risk premium channel, stock price movements following a monetary policy shock may be explained by the initial overreaction of investors and/or changes in investors' sentiment (Bernanke and Kuttner, 2005; Kurov, 2012).

French, 1995; 1996). Consequently, monetary policy decisions are likely to matter more for value firms as compared to firms with low E/P, BE/ME, C/P and D/P ratios, i.e. growth firms. As Lioui and Maio (2014) demonstrate, value stocks earn an additional risk premium relative to growth stocks with respect to future changes in interest rates. The lack of growth opportunities for value firms despite their stable cash flows make their stocks look more like long-term bonds.

Overall, there is a strong rationale why monetary policy actions could affect stock prices and reflect a type of a risk relevant for asset pricing. The following three sections summarise the existing empirical evidence with respect to monetary policy effects on stock prices in the US and discuss the bi-directional relationship between stock prices and monetary policy.

3.3 Monetary policy effects: Early findings

The empirical literature on the relationship between monetary policy actions and stock prices dates back to the 1970s.[122] Several early studies provide some evidence that increases in money stock lead to higher aggregate stock price level in the post-war period (Homa and Jaffee, 1971; Keran, 1971). On the other hand, others argue that causality between money and stock prices may be bi-directional (Cooper, 1974; Rogalski and Vinso, 1977). If the efficient market hypothesis holds, current stock prices reflect all available information and may help to forecast changes in money stock. Thus, only unexpected monetary policy actions should have an impact on asset prices. Subsequently, later empirical studies mainly use an event-study approach and focus on the effects of unexpected money supply changes. Their results show that unanticipated increases in the US money supply cause an immediate and significant decline in the stock market (Cornell, 1983; Pearce and Roley, 1983).[123]

[122] Sellin (2001) provides a detailed survey of early studies on the interaction between monetary policy and stock prices.

[123] Generally, monetary expansion is associated with increasing stock prices. This “puzzling” negative response can be explained in several ways. If higher than expected growth rate of money stock increases inflation expectations, stock prices may fall due to lower expected future real earnings. Also, market expectations about tighter future monetary policy following an unanticipated increase in money supply may lead to higher expected market interest rates lowering current stock prices. In addition, Cornell (1983) suggests the risk premium hypothesis: a positive surprise in money stock may indicate higher risk aversion among market participants and increased risk perception that causes stock prices to decline.

The discount and surcharge rates have also been used to measure monetary policy shifts.[124] For instance, Waud (1970) analyses the discount rate announcement effects over the period 1952 – 1967. The study shows that stock market returns are generally negative around the dates of discount rate increases, while discount rate cuts are associated with positive stock returns. Pearce and Roley (1985) find that the discount rate announcements have no significant impact on daily stock market returns prior to October 1979. On the other hand, they provide the evidence that the stock market responds negatively and significantly to discount rate and surcharge rate changes on the announcement days after October 1979. The study by Smirlock and Yawitz (1985) distinguishes between endogenous and exogenous discount rate changes, where the latter contain monetary policy-related information. They show that a stock price reaction to exogenous discount rate changes is negative and statistically significant in the post-October 1979 period. Likewise, Hardouvelis (1987) also reports a negative and significant effect of actual changes in discount and surcharge rates on stock returns after October 1979.

In the literature it has been recognised that the estimated asset price response to monetary policy actions may be biased due to the endogeneity issue (Smirlock and Yawitz, 1985; Cook and Hahn, 1989; Lee, 1992; Thorbecke and Alami, 1994; Thorbecke, 1997). Firstly, the causality may stem from the stock market to monetary policy (reverse causality). Forward-looking financial markets may provide monetary policymakers with useful information about the future path of economy and thus may be important to policy decision making (Bernanke and Gertler, 1999; Cecchetti et al. 2000; Baxa et al., 2013). Secondly, stock prices and monetary policymakers may simultaneously respond to other information, such as macroeconomic news. Based on the empirical approach employed to isolate truly exogenous unanticipated monetary policy changes, the related literature is split into two main strands: event studies and structural vector autoregression models (SVARs). The following sections review the empirical evidence with respect to these two methodologies.

3.4 Monetary policy effects: Event studies

[124] In the late 1980s, the federal funds rate targeting was initiated following the period of the borrowed reserves operating procedure (Strongin, 1995; Walsh, 2003). Thus, the majority of later empirical studies uses changes in the federal funds (target) rate to measure monetary policy shocks.

Using information in the *Wall Street Journal* on the days following a change in the federal funds target rate, Cook and Hahn (1989) construct the time series of target rate changes to measure exogenous monetary policy shocks and examine the response of market interest rates to these shocks.[125] Thorbecke and Alami (1994) employ this dataset by Cook and Hahn (1989) to investigate the US stock market response to monetary policy in the period 1974:9 - 1979:9. They find that target rate changes have a negative and significant effect on daily stock market returns. In the same spirit, Thorbecke (1997) uses several major newspapers and constructs the series of target rate changes for the period 1987:8 - 1994:12. The results show that the policy-induced increases in the funds target rate cause stock market returns to decline significantly. Nevertheless, it is unlikely that forward-looking financial markets respond to the policy actions that are expected. In order to distinguish between the expected and unexpected components of target rate changes, financial market or survey data may be used (Rudebusch 1998; Kuttner, 2001; Ehrmann and Fratzscher, 2004). For instance, Kuttner (2001) constructs daily and monthly measures of monetary policy surprises using the federal funds futures data to gauge market expectations about the federal funds rate. This approach has been widely employed in the related literature ever since.

Bernanke and Kuttner (2005) apply the technique developed by Kuttner (2001) and analyse the effect of monetary policy on the stock market over the period 1989:6 - 2002:12. Daily stock returns are regressed on the monetary policy surprises over the Federal Open Market Committee (FOMC) meeting days and the days of target rate changes. The empirical evidence is in favour of a negative and significant relationship between unexpected changes in the target rate and market returns. Similarly, Ehrmann and Fratzscher (2004) identify monetary policy surprises taking the difference between the FOMC announced target rate and the expected target rate based on the Reuters poll conducted prior to each FOMC meeting. With respect to the stock market, they find a negative and significant response of a daily return to unexpected monetary tightening over the period 1994:2 - 2003:1. In general, other studies also confirm the negative stock market’s response to positive target rate surprises (Guo, 2004; Basistha and Kurov, 2008; Jansen and Tsai, 2010; Kurov, 2010). In addition, Basistha and Kurov (2008) show that the effect of monetary policy on the stock market is significantly stronger during the periods of economic recession and tight credit market conditions. Another type of asymmetry in the monetary policy effect relates to the stock market conditions. In response to contractionary

[125] They argue that the reverse causality is not an issue due to delays in the policy implementation at that time.

monetary policy, the stock market declines significantly more in a bear market as compared to a bull market (Jansen and Tsai, 2010; Kurov, 2010).

With respect to individual stock and stock portfolio returns, the empirical evidence indicates highly heterogeneous monetary policy effects. For instance, Erhmann and Fratzscher (2004) analyse the five hundred individual stocks included in the S&P500 index. The results of the panel estimations show that industry- and firm-specific factors may help to explain the differences in stock price responses to monetary policy shocks over the period 1994:2 - 2003:1. The stocks of firms operating in cyclical and capital-intensive industries decline significantly more following monetary policy contraction than an average stock price. In addition, monetary policy has a significantly stronger effect on smaller and financially constrained firms (Erhmann and Fratzscher, 2004). Similarly, Basistha and Kurov (2008) examine the FOMC announcement effects on the S&P500 stocks over the period 1990:1 – 2004:12. They confirm that cyclical and capital intensive industries are more sensitive to monetary policy shocks. Furthermore, financially constrained firms experience sharper declines in stocks prices following monetary policy tightening than relatively unconstrained firms, especially in bad economic times. For the period 1990:1 – 2004:11, Kurov (2010) demonstrates that stocks with higher sensitivity to changes in the investor sentiment react more strongly to policy shocks in bear markets.

Stock returns on size- and book-to-market-sorted portfolios are analysed by Guo (2004) over two sample periods 1974:9 – 1979:9 and 1988:10 – 2000:2. In the earlier period that is associated with generally tight business conditions, smaller firms' stocks decline significantly more as compared to medium and large firms following an unexpected increase in the policy rate. Also, the effect of monetary policy is stronger for stocks with a high book-to-market ratio, i.e. value stocks. However, both the "size effect" and the "value effect" of monetary policy disappear in the later period that is associated with generally good business conditions.[126] On the other hand, Jansen and Tsai (2010) report that there is a more pronounced and significant decline in stock returns following a contractionary policy shock for smaller firms and firms in transportation, communication, services, manufacturing and retail trade industries in the period 1994:2 – 2005:12. In contrast, Cenesizoglu (2011) finds that larger stocks and growth stocks are more sensitive

[126] Throughout the text, the size effect refers to the differential impact of monetary policy on small stocks versus large stocks, with small stocks being more responsive. Similarly, the value effect denotes the differential impact of monetary policy on the stock returns of portfolios formed on the basis of value characteristic proxies, with value stocks being more responsive.

to monetary policy shocks during 1989:6 - 2009:12. On the other hand, the difference in the response coefficients is not statistically significant.[127]

The majority of event studies uses daily data that may introduce a modest bias in the estimated stock price response to monetary policy shocks (Rigobon and Sack, 2004). Since high-frequency data mitigates the problems of reverse causality and omitted variables bias, researchers have turned to intraday data within the event-study framework (Gurkaynak, Sack, and Swanson, 2005; Ammer, Vega and Wongswan, 2010; Rosa, 2011).[128] Nevertheless, the results of daily event studies are largely confirmed using high-frequency data.

Overall, the event-study literature finds that monetary policy has a negative and significant impact on stock prices. Generally, small and financially constrained firms are more exposed to unexpected changes in monetary policy stance; however, more recent empirical evidence of the size and value effects with respect to a monetary policy shock is somewhat weaker and mixed.

3.5 Monetary policy effects: Structural VARs

As the alternative to large-scale structural macroeconomic models widely used in the 1970s, Sims (1980) proposed a vector autoregression (VAR) model. The *k*-equation model defines each of *k* endogenous variables as a linear function of its own lags and the lagged values of the remaining *k-1* variables. The model can then be estimated equation by equation using the standard ordinary least squares (OLS) method. VARs are valuable tools to describe data and to produce forecasts. Nevertheless, such a reduced-form model says nothing about the structural interpretation of macroeconomic relationships in the defined system (Stock and Watson, 2001). The residuals are correlated across the equations since endogenous variables are correlated with each other. In order to recover uncorrelated fundamental economic shocks, one needs to disentangle the innovations of a structural VAR (SVAR) model using its reduced-form residuals. This requires some economic theory to restrict contemporaneous relationships in the system (Sims, 1986; Stock and Watson,

[127] Some academics argue that the stock market anomalies may be a time-varying phenomena (Horowitz, Loughran and Savin, 2000; Hahn, O'Neill, and Reyes, 2004; van Dijk, 2011). This could possibly explain the mixed evidence of a significant differential monetary policy effect on stock returns.

[128] On the other hand, Thornton (2014) discusses two issues related to the usage of intraday data. Firstly, financial markets may over-react to policy actions leading to some noise. Secondly, not all policy decisions used to be announced. For instance, the FOMC only started announcing its decisions since February 1994.

2001).[129] Then, the identified SVAR model allows generating the dynamic responses of endogenous variables to each fundamental shock, known as impulse response functions (IRFs). It may also be used for the forecast error variance decompositions, historical decompositions and simulations. Overall, SVARs provide a powerful and popular tool for the empirical macroeconomic analysis and are extensively applied in empirical work (Kilian, 2011).

The seminal work by Bernanke and Blinder (1992) shows that the innovations from the federal funds rate equation in a structural VAR model could be interpreted as monetary policy shocks. Subsequently, it since has become a common practice to use SVAR models to estimate both the macroeconomic effects of monetary policy (Christiano, Eichenbaum and Evans, 1996; Kim, 2001; Leeper and Zha, 2003; Mackowiak, 2007; Barakchian and Crowe, 2013; Doehr and Martinez-Garcia, 2015) as well as the financial effects (Thorbecke, 1997; Park and Ratti, 2000; Rapach, 2001; Vargas-Silva, 2008; Bjornland and Leitemo, 2009; Hammoudeh, Nguyen and Sousa, 2015).[130] Nevertheless, there has been a great deal of debate about the appropriateness of identifying restrictions commonly used to disentangle structural monetary policy shocks in SVAR models (Stock and Watson, 2001; Kilian, 2011). Over time, new approaches and strategies have been developed in the literature and the innovations to this respect still continue.

3.5.1 Recursive (Cholesky) identification

The recursive identification scheme rests on the recursive ordering of endogenous variables implying specific causal relationships and it is based on the Cholesky decomposition of the variance-covariance matrix of reduced-form errors (Christiano, Eichenbaum and Evans, 1998; 2005). Typically, some contemporaneous coefficients are set to zero so that the matrix of contemporaneous coefficients is a lower triangular matrix. For instance, Christiano, Eichenbaum and Evans (1998) separate a state vector of endogenous variables into three blocks. The first block variables have a contemporaneous effect on a monetary policy instrument, but they respond only with a lag to the remaining variables in the system. The policy instrument variable is placed between the first and the third blocks. This implies that policymakers observe and respond contemporaneously to the variables in the first block, but they react only with a lag to those variables in the third

[129] Sims (1980) recovers structural shocks assuming a triangular matrix of contemporaneous response coefficients. In this way, the first variable in the system is only explained by the lagged values of all endogenous variables, while the last variable is also influenced contemporaneously by all other variables.

[130] Generally, it is only one structural shock identified in such studies, i.e. the monetary policy shock. Hence, these models are sometimes referred to as semi-structural VARs (Killian, 2011).

block. Those variables ordered after the policy block are assumed to be contemporaneously affected by all preceding variables (CEE, 1998).

Numerous empirical studies have used this identification scheme to estimate the effects of monetary policy on stock prices where stock prices are ordered the last in the VAR state vector. It is equivalent to the assumption that policymakers do not take into account current stock prices when making policy decisions, although stock prices adjust to monetary policy news immediately. For instance, Thorbecke (1997) investigates monetary policy effects on twenty two industry and ten size-sorted stock portfolios in the period 1967:1 – 1990:12. The initial-period responses show that an unexpected one-standard-deviation increase in the federal funds rate has a negative and generally significant impact on monthly stock returns of about 0.8% on average. The response coefficients are more negative for smaller stocks as compared to larger stocks offering support for the credit channel of monetary policy transmission. Nevertheless, there is no monotonic decline in the magnitude of the response parameters across size-sorted portfolios. In addition, Thorbecke (1997) also provides the evidence of a heterogeneous monetary policy impact on returns across industry stock portfolios.

In addition to stock prices, Cheng and Jin (2013) also include a term spread and house prices in a SVAR model. All three financial variables are assumed to be influenced contemporaneously by policy shocks with stock prices being the most responsive variable. Similarly, the findings indicate that monetary policy contraction has a negative and significant impact on the stock market for the period 1979:Q3 - 2006:Q1. A one-standard-deviation positive shock decreases stock market returns by 1.5%. Moreover, the stock market appears to have an indirect effect on monetary policy through its impact on inflation and output. A positive stock price shock has a delayed, positive and significant effect on the federal funds rate.

The empirical analysis in Sousa (2014) uses the Bayesian methods to estimate a SVAR model and to pin down the impact of US monetary policy on housing and financial wealth as well as their components in the period 1947:Q1 - 2008:Q4. Overall, the results imply that there is a negative and significant effect on asset wealth. With respect to financial wealth, an unexpected increase in the federal funds rate results in a relatively short-lived but statistically significant fall in net financial wealth. Also, the stock market declines significantly in response to unexpected monetary policy tightening (Sousa, 2014).

With respect to time-varying monetary policy effects, the study by Chang, Chen and Leung (2011) employ a regime-switching SVAR model to analyse US monetary policy effects on various asset prices in the period 1975:Q1 - 2008:Q1. They identify two regimes: a high-volatility regime in the late 1970s and early to the mid-1980s, while the period starting in the mid-1980s is identified as a low-volatility regime. The findings show that, following a contractionary monetary policy shock, stock market returns decline and this decline is greater in magnitude in the low-volatility regime as compared to the high-volatility regime. Park and Ratti (2000) estimate a rolling VAR model and investigate the interactions between monetary policy and expected real stock market returns in the US. Over the period 1973:1 - 1998:3, they find a significant stock market reaction to monetary policy shifts. In the period since the early 1980s, contractionary monetary policy shocks have somewhat less negative effects on real expected stock returns as compared to the earlier period. In the similar spirit, Gali and Gambetti (2014) investigate whether monetary policy shocks in the US have any effects on stock market bubbles in 1960:Q1 - 2011:Q4. They estimate a time-varying parameter SVAR using the Bayesian methods. The results show that the fundamental component of stock prices always declines in response to monetary policy contraction. The negative response of the fundamental component remains stable over time. On the other hand, the response of stock prices is time-varying. Initially, the stock market declines quite substantially in response to an unexpected hike in an interest rate. However, the decline seems to be much more persistent during the 1970s. starting in the early 1980s, the initial drop reverses quickly with stock prices overshooting the initial level. Gali and Gambetti (2014) argue that this finding is consistent with the theory of rational asset price bubbles. In other words, higher interest rates lead to greater expected stock price growth in the presence of a relatively large bubble (Gali, 2014).

3.5.2 Non-recursive identification

The recursive identification strategy has received substantial criticism in the literature due to the lack of economic reasoning and controversial short-run zero restrictions (Carlstrom, Fuerst and Paustian, 2009; Castelnuovo, 2013). An alternative approach is to employ the non-recursive identification as advocated by Leeper, Sims and Zha (1996) and Sims and Zha (2006). Within the framework of Sims and Zha (2006), monetary policy is measured by total reserves and the funds rate. Monetary policy is assumed not to respond contemporaneously to price level and output due to the fact that there is no contemporaneous macroeconomic data available to policymakers at the time of decision making. Consequently, the interest rate rule includes the contemporaneous values of the producers’ price index for intermediate goods and total reserves alongside the lags of all variables. Monetary policy has only a lagged impact on private sector variables, such as the producers' price index for intermediate materials, real output, the gross national product deflator, average hourly earnings and bankruptcy filings. Within the block of these sluggish variables, the recursive ordering applies. Meanwhile, the money demand function links contemporaneously total reserves to real output, the price deflator and a short-term interest rate. Finally, the producers' price index for intermediate goods is influenced by all variables contemporaneously providing an indirect contemporaneous link between monetary policy and the private sector variables (Sims and Zha, 2006).

Similar strategies have been since applied in the SVAR-based analysis of monetary policy including asset prices. For instance, Li, Iscan and Xu (2010) use a modified version of the non-recursive identification to compare the real stock price response to monetary policy shocks in the US and Canada during the period 1988:1 - 2003:12. The interest rate rule includes the contemporaneous values of money stock and the lagged values of real output, aggregate price level, money stock, the funds rate, oil and stock prices. The stock market is allowed to respond without a delay to all information in the system. Nevertheless, this identification approach does not allow for the simultaneous interaction between the stock market and the policy interest rate. Li, Iscan and Xu (2010) find that an unexpected increase in the funds rate of 25 basis points leads to an instant decline of 0.55% in the stock market index with the effect becoming statistically significant after several periods.

Some analysis of the interdependence between the Fed's policy and the stock market is presented by Chatziantoniou, Duffy and Filis (2013). In their identification scheme, money supply is contemporaneously linked to inflation, output, and government expenditure. In addition, a short-term interbank interest rate responds to the current values of money supply, aggregate stock prices, global economic conditions and government expenditure. Finally, the stock market is influenced contemporaneously by all variables, but stock price developments have an immediate effect only on the interbank interest rate. For the period 1991:Q1 - 2010:Q4, the evidence indicates the bi-directional relationship between the interbank interest rate and the stock market. An exogenous increase in money supply drives down the short-term interest rate and increases stock prices. Meanwhile, a positive stock price shock induces an increase in the interbank interest rate (Chatziantoniou, Duffy and Filis, 2010).

3.5.3 Generalised impulse response functions

While the non-recursive scheme may seem attractive, there are some costs associated with it. The more complex identification scheme requires a broader set of contemporaneous economic relationships to be defined. Moreover, some of assumptions about the short-run dynamics are just as debatable as those imposed in the recursive identification. The alternative strategy to generate impulse responses from reduced-form VARs is proposed by Pesaran and Shin (1998). The suggested generalised impulse response functions do not require the orthogonalisation of reduced-form residuals and short-run restrictions on contemporaneous relationships between endogenous variables. Moreover, the estimation results are independent of the ordering of variables.

Ewing, Forbes and Payne (2003) use this approach to investigate how US monetary policy shocks affect returns on the sector-specific stock market indices for the period 1988:1 - 1997:7. The following sectors are investigated: financials, capital goods, industrials, transportation and utilities. They find that returns on all indices decline significantly in response to an unexpected increase in the funds rate. The largest effect materialises in the sectors of financials and capital goods. Nevertheless, the impact of monetary policy shock generally dissipates within two months (Ewing, Forbes, and Payne, 2003).

The recent study by Kontonikas and Kostakis (2013) also takes an advantage of the generalised impulse response approach. Essentially, they extend the analysis by Thorbecke (1997) and estimate the response of stock returns on portfolios sorted by various characteristics to monetary policy shocks. The results for the sample period 1967:1 – 2007:12 provide the evidence in favour of the credit and risk premium channels of monetary policy transmission. Generally, stock returns on all portfolios decline significantly in response to monetary policy tightening. Moreover, value stocks decrease significantly more than growth stocks, whilst small stocks are also more negatively affected than large stocks. Furthermore, the portfolios of stocks that performed poorly in the past are also somewhat more responsive to policy shocks as compared to the past winner stocks. However, the sub-sample analysis reveals that the full-sample results are mostly driven by the pre-1983 period. There is no evidence of the significant monetary policy impact on stock returns and no evidence for the differential impact of monetary policy shock on stock returns across portfolios in the post-1983 period (Kontonikas and Kostakis, 2013).

3.5.4 Long-run restrictions

The recursive and non-recursive identification schemes use short-run restrictions on the matrix of contemporaneous parameters. In contrast, Blanchard and Quah (1989) introduced long-run restrictions in order to identify structural shocks in VAR models. Long-run restrictions have since been applied in the structural VAR models that estimate an asset price response to monetary policy shocks. Generally, the long-run monetary neutrality is assumed for such restrictions to hold, i.e. monetary policy shocks have no long-run effects on real variables at infinite horizons. This assumption is in line with many theoretical economic models. Nevertheless, Faust and Leeper (1997) argue that using long-run restrictions may produce unreliable results due to the fact that long-run relationships may not be captured well in finite sample periods. Instead of applying such restrictions for the infinite horizon, they suggest that either finite-horizon long-run or standard short-run restrictions should be preferred.[131]

For the US, Lastrapes (1998) measures monetary policy shocks using money supply and assumes that exogenous shifts in nominal money supply have no permanent impact on the levels of real macroeconomic variables, including real stock prices. Thus, short-run dynamics among variables in the system remain unrestricted. For the period 1960:3 – 1993:12, the evidence indicates the stock market liquidity effect. Following a positive permanent 1% shock in nominal money supply, the real stock market increases significantly by 2.4%. Nevertheless, this study does not analyse a systematic monetary policy response to developments in the stock market. The related study by Rapach (2001) takes a similar approach in order to examine real US stock price response to money supply innovations during 1959:Q3 – 1999:Q1. In line with Lastrapes (1998), a positive shock in money supply leads to a significant increase in real stock market prices. Moreover, Rapach (2001) demonstrates that the 3-month Treasury bill rate increases significantly in response to a positive real stock price shock. This indicates that the Federal Reserve may act in order to curb stock prices by raising interest rates. Similarly, Crowder (2006) estimates a daily SVAR model containing only the federal funds rate and stock market returns for the period 1970:2 – 2003:6. The long-run restriction implies that a stock price shock has no effect on the federal funds rate in the long run. However, his findings are at odds with the majority of studies. Firstly, equity returns significantly increase in response to monetary policy contraction. Secondly, the funds rate decreases significantly in response to unexpectedly higher stock prices.

[131] Lastrapes (1998) addresses this critique by assuming different finite horizons of 1, 6 and 48 months when imposing long-run restrictions. The results show that impulse response functions using the 48-month horizon are almost identical to those based on the infinite-horizon long-run restrictions.

Bjornland and Leitemo (2009) suggest the strategy that does not solely rely on long-run restrictions over the infinite horizon. Structural shocks are identified using the combination of short-run and long-run restrictions that allows for the contemporaneous interaction between the stock market and the monetary policy rate. With respect to standard macroeconomic variables, the recursive identification is applied. The long-run restriction implies that monetary policy does not have an impact on real stock prices in the long run. In this way, both the federal funds rate and real stock returns react contemporaneously to all information in the SVAR. They find the evidence of the strong and significant interdependence between the funds rate and real stock returns in the period 1984:6 – 2002:12 (Bjornland and Leitemo, 2009). An unexpected increase in the federal funds rate by 1 percentage point is associated with an instant drop in real stock prices of around 9%. The effect is also statistically significant and quite persistent. At the same time, the funds rate increases by about 4 basis points in response to a 1% positive real stock price shock and continues upwards for about a year. The response is also statistically significant.

Bjornland and Jacobsen (2013) extend the model of Bjornland and Leitemo (2009) to include both stock prices and house prices in the analysis for the US. The findings for the period 1983:Q1 - 2010:Q1 are consistent with those reported in Bjornland and Leitemo (2009). A contractionary monetary policy shock is associated with an immediate decrease in the stock market of about 10%. The response is significant for several periods; nevertheless, it is quickly reversed. Meanwhile, the federal funds rate increases following a positive shock to stock market prices. This implies the simultaneous relationship between monetary policy and the stock market.

Within a similar framework, Laopodis (2013) analyses the dynamic relationship between the US stock market and monetary policy over the chairmanship of Burns, Volcker and Greenspan. Generally, stock prices respond negatively to a contractionary monetary policy shock across the three eras, but the response is time-varying. In addition, the funds rate reaction to stock price shocks also appears to depend on the sample period. Overall, Laopodis (2013) suggests that there is no clear and consistent dynamic relationship between the policy rate and the stock market.

3.5.5 Sign restrictions

In light of criticism towards traditional identification schemes, sign restrictions have become increasingly popular in recent years (Faust, 1998; Canova and de Nicolo, 2002; Uhlig, 2005). Structural shocks from the estimated VARs are recovered by imposing signs on the contemporaneous response parameters of endogenous variables based on some economic theory. The identification scheme based on sign restrictions has been also applied in the empirical work investigating conventional monetary policy effects on asset prices (Vargas-Silva, 2008; Bjornland and Jacobsen, 2013) as well as the effects of quantitative easing policies (Gambacorta, Hofmann and Peersman, 2014).

For instance, Bjornland and Jacobsen (2013) use sign restrictions as an alternative identification scheme in the robustness analysis of their study. The baseline specification is amended slightly and the state vector only includes output, inflation, the federal funds rate and real stock returns. In addition to the recursive restrictions for the first three variables, one sign restriction is imposed. It implies a non-positive initial stock price reaction in response to a contractionary monetary policy shock. The findings indicate a negative and statistically significant reaction of stock prices to an unexpected increase in the funds rate for several periods following the shock. Moreover, the policy rate also increases significantly in response to a positive stock price shock.

3.5.6 Heteroscedasticity-based identification

Researchers who use sign restrictions do not have to rely on other types of restrictions that may seem ad hoc or unrealistic. Nevertheless, there are several pitfalls associated with the usage of sign restrictions (Fry and Pagan, 2011; Kilian, 2011). This identification approach is rather agnostic and there is no unique point estimate of impulse response functions. Due to the lack of specified information to discriminate between the shocks, it is also likely that there is more than one structural shock of the same kind identified (Fry and Pagan, 2011). Given the shortcomings in many identification strategies, another strand of the empirical literature solves the identification problem in SVAR models by taking the advantage of heteroscedasticity present in the data (Lanne and Lutkepohl, 2008).

Using daily data, Rigobon and Sack (2004) analyse the impact of monetary policy on the US stock market in the period 1994:1 – 2001:11. Structural policy shocks are identified assuming that the variance of monetary policy shocks is greater relative to asset price shocks on the days of policy meetings. The results show that a positive shock to a short-term interest rate has a negative and statistically significant impact on stock prices. In the related study by Rigobon and Sack (2003), the identification of stock price shocks rests upon the observed shifts in the covariance matrix of reduced-form residuals and the assumption of homoscedastic monetary policy shocks. The evidence indicates that US monetary policy does react to stock prices in the period 1985:3 – 1999:12. Thus, both studies together provide the support for the interdependence between monetary policy and developments in the US stock market.

Similarly, Bouakez, Essid and Normandin (2013) use the heteroscedasticity-based approach with monthly data and revisit the simultaneity between stock returns and US monetary policy in the single study for the period 1982:11 – 2007:11.[132] They show that stock returns are not significantly affected by monetary policy shifts and the sign of the response coefficient is counter-intuitive, i.e. stock returns increase in response to monetary policy tightening. Moreover, stock prices do not seem to contain any relevant information to policymakers beyond their impact on aggregate economy. They also briefly analyse stock returns on portfolios formed on the industry, firm's size and BE/ME ratio. However, they do not report these results and the discussion of the results is not well-developed.

The recent study by Lutkepohl and Netsunajev (2015) takes the heteroscedasticity-based approach another step further. The proposed model determines shifts in the volatilities of shocks using a smooth transition function instead of assuming exogenous changes in the variance-covariance matrix of residuals.[133] With respect to stock prices, the results are in line with the existing literature and show that stock prices decline following an exogenous increase in the funds rate.

3.5.7 Other identification strategies

Several other strategies have been proposed in the literature to identify monetary policy shocks.[134] For instance, Bernanke and Mihov (1998) suggest a VAR-based methodology that does not assume a priori the monetary policy instrument but rather derives it by estimating the model of central bank's operating procedures. Other studies use financial market data to identify policy shocks outside a VAR model (Barakchian and Crowe, 2013) or, alternatively, a narrative approach to deduce exogenous policy changes (Romer and Romer, 2004; Bluedorn and Bowdler, 2011). The generated series of shocks are then used to estimate the model. Some others combine high and low frequency data.

[132] Their model combines the methodology of Bernanke and Mihov (1998) with the heteroscedasticity-based approach. They show that the data rejects both the recursive identification (Thorbecke, 1997) as well as that applied by Bjornland and Leitemo (2009). However, the model used by Bouakez, Essid and Normandin (2013) is not directly comparable to the model used in Bjornland and Leitemo (2009).

[133] While this study also rejects the identifying restrictions as in Bjornland and Leitemo (2009), they apply restrictions over the different sample period (1970:1 - 2007:6) than in the original study.

[134] While this chapter attempts to review thoroughly the major developments in the SVAR-based analysis, the list of alternative approaches is by no means exhaustive.

For instance, D'Amico and Farka (2011) use intraday data on the federal funds futures contract rates to identify structural policy shifts on the days of policy announcements. Then, stock returns are regressed on these shocks to estimate the response coefficient of stock prices that is then applied as a restriction in a monthly VAR model. Their results are supportive of the strong interdependence between the US stock market and monetary policy for the period 1994:1 – 2006:9. Finally, Tsai (2011) measures monetary policy shocks based on an autoregressive conditional hazard VAR model and distinguishes two sources of positive monetary policy surprises. The first type of a surprise refers to an unexpected increase in the federal funds rate when it is expected to remain unchanged. The second type of a monetary policy surprise captures an unexpected increase in the funds rate when it is expected to be decreased but is instead kept constant. The results indicate that the first-type policy surprises have a stronger negative effect on stock returns with respect to market level and size-sorted portfolios. The empirical evidence is in favour of the size effect of monetary policy for the period 1984:3 – 2008:9 (Tsai, 2011).

3.5.8 Developments in other dimensions

While there has been a great variety of innovations with respect to the identification of structural VAR models, this is not the only dimension of developments in the literature. In order to account for potential structural changes in the dynamics of relationships among variables over time, it becomes increasingly popular to employ time-varying structural VAR models for monetary policy analysis (Primiceri, 2005; Canova and Gambetti, 2007). Standard VAR models typically contain a relatively small number of variables as compared to the actual information set available to monetary policymakers at the time of decision making. Thus, a growing number of studies rely on the methods that allow dealing with large datasets within the SVAR framework. For instance, Bernanke, Boivin and Eliasz (2005) combine a factor analysis with a standard SVAR approach (FAVAR). This way, a large set of information is summarised by several factors. Alternatively, one may choose to use a large-scale Bayesian VAR model to circumvent the issue of omitted variable bias (Banbura, Giannone and Reichlin, 2010). Finally, panel VARs are helpful in accounting for common dynamic relationships across countries and improving the accuracy of a structural analysis (Gambacorta, Hofmann and Peersman, 2014).

3.6 Methodology

3.6.1 Motivation

Generally, SVAR-based studies tend to focus on the market level when analysing monetary policy effects on stock prices. Several existing articles that do examine stock portfolios, i.e. the size and value effects of monetary policy, typically employ an identification scheme without fully taking into account the simultaneous interaction between monetary policy and stock prices (Thorbecke, 1997; Tsai, 2011; Kontonikas and Kostakis, 2013). To fill the gap in the literature, this chapter employs a constant-parameter structural VAR model for the US in the spirit of Bjornland and Leitemo (2009) to thoroughly analyse the bi-directional causality between monetary policy and stock prices at both market and portfolio levels. The main advantage of the original specification is that it fully takes into account the potentially simultaneous interaction between monetary policy and real stock prices. This identification of monetary policy shocks is achieved through a combination of standard short-run restrictions and one long-run restriction on the basis of long-run monetary policy neutrality. As monetary policy instrument is assumed to have no long-run effects on real stock prices, the contemporaneous relationship between the two variables remains intact.[135]

Structural VAR models used for the analysis of monetary policy effects are often criticised for the omission of potentially important informational variables and the inadequate description of the monetary policy rule. The price puzzle often reported in the empirical work is considered to indicate the problem of model misspecification and omitted variables (Rusnak, Havranek and Horvath, 2013).[136] Consequently, it has become a standard practice in the literature to include a commodity price index to help forecast inflationary pressures and to account for potentially omitted information (Sims, 1992; Thorbecke, 1997, Bjornland and Leitemo, 2009). However, it has been shown that the forecasting power of a variable may not be related to its ability to eliminate the price puzzle (Hanson, 2004). In response, several alternatives how to eliminate the puzzle have been proposed. For instance, Giordani (2004) argues in favour of the output gap as a measure of economic activity. Krusec (2010) demonstrates that the price puzzle can be resolved by imposing long-run identifying restrictions. Furthermore, Rusnak, Havranek,

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[135] The combination of short- and long-run restrictions also addresses the criticism towards the identification schemes solely based upon long-run restrictions (Faust and Leeper, 1997; Abouwafia and Chambers, 2015).

[136] The price puzzle refers to a positive response of prices following a contractionary monetary policy shock. According to Sims (1992), the price puzzle may be generated if some information that policymakers have about inflationary pressures is not included in the estimated model. For instance, following pre-emptive monetary policy tightening due to higher expected future inflation, the price level will rise, albeit, perhaps, to a smaller degree than otherwise.

and Horvath (2013) argue that the puzzle appears due to the model misspecification. The model may be improved by the inclusion of the commodity prices and output gap, the application of non-recursive identification scheme and Bayesian estimation methods.

Some of the proposed techniques have been successful in addressing the price puzzle to some extent. Nevertheless, most of them still fail to specify appropriately the reaction function of a central bank that requires a large set of information available to monetary policymakers with forward-looking elements. To this regard, Brissimis and Magginas (2006) propose the alternative that solves the price puzzle and produces a sharper measure of monetary policy shocks within a simple SVAR model. They augment an otherwise standard VAR specification for the US by including two forward-looking variables. The first variable is the expected level of the policy interest rate as inferred from the federal funds futures contract rate. The second variable represents near-future expectations about economic developments and is measured by a leading indicator of economic activity. This approach controls more effectively for the information set that central bank uses for policy decision making. The amount of contemporaneously available information in the VAR system is increased considerably without a substantial reduction in the degrees of freedom or explosion in the dimension of the model (Brissimis and Magginas, 2006).

The first part of the empirical analysis in this chapter is based on the original SVAR model as defined in Bjornland and Leitemo (2009). Motivated by the above discussion, the original specification of the SVAR is augmented for the main analysis according to the recommendations by Brissimis and Magginas (2006).

3.6.2 Structural VAR

The p-order n-variable structural autoregressive model may be written in the following form (ignoring deterministic terms for notational convenience):

$$B_0 Z_t = B_1 Z_{t-1} + B_2 Z_{t-2} + ... + B_p Z_{t-p} + \varepsilon_t \tag{3.1}$$

where Z_t is the (n x 1) vector containing n endogenous variables, B_0 is the (n x n) matrix of contemporaneous coefficients, B_i is the (n x n) matrix of lag coefficients, for i = 1, 2,…, p, and ε_t denotes the (n x 1) vector of serially uncorrelated structural innovations with a zero mean and the variance-covariance matrix $\Sigma_\varepsilon = E(\varepsilon_t \varepsilon_t')$, i.e. ε_t is the vector of structural shocks (Kilian, 2011).[137] The model in Equation (3.1) can be expressed in a more compact form using a matrix polynomial in the lag operator L:

$$B(L)Z_t = \varepsilon_t \tag{3.2}$$

where $B(L) = B_0 - B_1L - B_2L^2 - \dots - B_pL^p$.

In Equations (3.1) and (3.2) some endogenous variables are allowed to interact contemporaneously. Consequently, the above model cannot be estimated using the standard OLS. Therefore, the structural VAR model has to be transformed into a reduced form to allow its estimation by the OLS. In order to derive the reduced-form representation of the model in Equation (3.1), both sides are pre-multiplied by the matrix B_0^{-1}:

$$B_0^{-1}B_0Z_t = B_0^{-1}B_1Z_{t-1} + B_0^{-1}B_2Z_{t-2} + \dots + B_0^{-1}B_pZ_{t-p} + B_0^{-1}\varepsilon_t \tag{3.3}$$

$$Z_t = A_1Z_{t-1} + A_2Z_{t-2} + \dots + A_pZ_{t-p} + w_t \tag{3.4}$$

$$A(L)Z_t = w_t \tag{3.5}$$

where $A_i = B_0^{-1}B_i$, for $i = 1, 2, \dots, p$, represents the reduced-form parameters and the reduced-form residuals denoted by $w_t = B_0^{-1}\varepsilon_t$ are serially uncorrelated, with a zero mean and the constant variance-covariance matrix Σ_w. Hence, w_t is a linear combination of the structural innovations ε_t. Also, $A(L) = I - A_1L - A_2L^2 - \dots - A_pL^p$.

After the reduced-form VAR model is estimated using the OLS, the structural form of the model can be recovered in order to learn about the responses of endogenous variables to the identified structural shocks. From the above, it is clear that the knowledge of B_0^{-1} matrix allows the calculation of the structural shocks using $\varepsilon_t = B_0w_t$, and the reconstruction of structural parameters using the relationship $B_i = B_0A_i$. The OLS estimation of Equation (3.5) provides n^2p coefficients in $A(L)$ and another $(n^2 + n)/2$ distinct parameters are obtained from the estimated Σ_w since it is a symmetric matrix. Nevertheless, there are $n^2 + n^2p$ free structural parameters in $B(L)$ and another $(n^2 + n)/2$ unique structural elements in the variance-covariance matrix of the structural residuals. As a result, there is the total of n^2 unknown variables in the structural system and further restrictions are needed in the system (Enders, 2015).

Given that $w_t = B_0^{-1}\varepsilon_t$, the variance-covariance matrix of the reduced-form residuals may be written as follows:[138]

[137] Note that B_0 is an invertible, square matrix and Σ_ε is a positive definite matrix (CEE, 1998).

$$\Sigma_w = B_0^{-1}\varepsilon_t (B_0^{-1}\varepsilon_t)' = B_0^{-1} E\left(\varepsilon_t \varepsilon_t'\right) B_0^{-1\prime} = B_0^{-1}\Sigma_\varepsilon B_0^{-1\prime} \tag{3.6}$$

In the SVAR literature it is common to assume that structural innovations are mutually uncorrelated implying that Σ_ε is a diagonal matrix. Furthermore, structural shocks typically are normalised to have a unit variance.[139] Thus, the variance-covariance matrix Σ_ε is the identity matrix *I* and Equation (3.6) may be re-written in the following form:

$$\Sigma_w = B_0^{-1} B_0^{-1\prime} \tag{3.7}$$

The orthogonality assumption is equivalent to $(n^2 - n)/2$ restrictions, while the normalisation of the variance of the structural shocks provides additional *n* restrictions. The OLS estimation of the reduced-form VAR representation provides the estimates of the left-hand side term in Equation (3.7). If the number of parameters in B_0^{-1} is not larger than the number of equations, i.e. unique elements $(n^2 + n)/2$ in Σ_w, the model can be identified. Since there are n^2 unknown elements in B_0^{-1}, after obtaining the estimate of Σ_w, a total of $(n^2 - n)/2$ additional restrictions is required to be imposed on B_0^{-1}, or equivalently, B_0 (Killian, 2011; Enders, 2015). Usually, these restrictions are in the form of short-run zero restrictions, for instance, the Cholesky factorization of Σ_w, or sign restrictions, and etc.

The alternative approach is to apply restrictions on long-run relationships among endogenous variables (Blanchard and Quah, 1989) or use them in a combination with short-run restrictions (Gali, 1992; Bjornland and Leitemo, 2009). With respect to long-run restrictions, consider an endogenous variable that contains a unit root but is difference-stationary and enters a VAR model in a differenced form. The long-run restriction of no permanent effect of a structural shock on the *level* of this variable implies that the cumulated impact of the structural shock on its differences must be equal to zero.

In order to implement long-run restrictions, a VAR model must be expressed in a vector moving average (VMA) form.[140] The VMA representation of the SVAR model in Equation (3.2) is:

$$Z_t = B(L)^{-1}\varepsilon_t = C(L)\varepsilon_t \tag{3.8}$$

[138] Note that the formula uses the property of transpose matrices (AB)′ = B′A′. This property implies that the transpose of a product of matrices is equal to the product of their transposes in a reverse order.

[139] In this case, the diagonal elements of B_0 are not restricted. Alternatively, one could impose restrictions so that the diagonal elements of B_0 are equal to unity, while leaving unrestricted the diagonal elements of Σ_ε. See Killian (2011) for more details.

[140] Such representation allows tracing out the response of a variable to various shocks over time, i.e. it produces impulse response functions.

where $C(L)$ denotes the (n x n) matrix of polynomial lags $C(L) = [C_{ij}(L)]$, for $i, j = 1,\ldots, n$. This way, each endogenous variable in the system is expressed in terms of the current and past structural shocks. An individual coefficient $c_{ij}(k)$ of a polynomial $C_{ij}(L) = \sum_{k=0}^{\infty} c_{ij}(k)$ denotes the response of a variable i to a structural shock in a j variable after k periods. The long-run restriction of no permanent effect of a variable j on the *level* of a difference-stationary variable i that enters VARs in a stationary form implies that the infinite cumulative effect of the structural shock ε_j on Δi must be equal to zero, i.e. $\sum_{k=0}^{\infty} c_{ij}(k) = 0$ (Blanchard and Quah, 1989).

The coefficients in $C(L)$ must be recovered from the estimated VAR model. Provided that the model satisfies the stability condition and is invertible, the corresponding reduced-form VMA representation is:

$$Z_t = A(L)^{-1} w_t = D(L) w_t \tag{3.9}$$

The relationship between the structural and reduced-form residuals is defined as:

$$w_t = B_0^{-1} \varepsilon_t \tag{3.10}$$

From Equations (3.8) - (3.10) it follows that:

$$Z_t = C(L)\varepsilon_t = D(L) w_t = D(L) B_0^{-1} \varepsilon_t \tag{3.11}$$

$$C(L) = D(L) B_0^{-1} \tag{3.12}$$

For $L = 1$:

$$C(1) = D(1) B_0^{-1} \tag{3.13}$$

where the matrix $C(1)$ represents the long-run responses of endogenous variables to the structural shocks and $D(1) = A(1)^{-1}$, $A(1) = I - A_1 - \ldots - A_p$.

The OLS estimation of the reduced-form VAR model gives the parameters of $A(1)$ matrix polynomial that is then inverted to obtain the estimate of $D(1)$. If the structural shocks are assumed to be orthogonal and normalised with a unit variance, the reduced-form covariance matrix is:

$$\Sigma_w = B_0^{-1} B_0^{-1\prime} \tag{3.14}$$

The structural parameters can be identified if there are enough restrictions placed on the contemporaneous matrix of the structural coefficients (alternatively, B_0^{-1}) and/or on the matrix of the long-run responses of variables to the structural shocks $C(1)$.

3.6.3 Baseline model: specification and identification

The study by Bjornland and Leitemo (2009) is key for the empirical analysis in this chapter. The estimated models are specified on the basis of the original SVAR as determined in Bjornland and Leitemo (2009). The starting point is to replicate their study for the original sample period. Next, the sample period is extended until the global financial crisis, keeping the specification of the model identical, i.e. the baseline model of this chapter. Finally, the specification is augmented by replacing the commodity price inflation variable with two forward-looking variables. The state variables are defined in line with Bjornland and Leitemo (2009). This allows to compare the results reported in this chapter with the original study and to evaluate the augmented model.

This section presents the baseline specification. The baseline SVAR model contains the output gap (gap_t), the first difference in the annual consumer price inflation ($\Delta\pi_t^a$), the annual commodity price inflation ($\pi_t^{comp,a}$), monthly real stock market returns (Δsp_t) and the monetary policy interest rate denoted by the federal funds rate (i_t). Note that this chapter defines inflation as the annual change in the price level since the monetary policy target for inflation is typically expressed in terms of annual inflation. On the other hand, inflation variable in Chapter 2 refers to a monthly change in the price level. The series of annual consumer price inflation is differenced to stationarity for the sample period considered. This approach is also taken by Bjornland and Leitemo (2009). Table A3.1 in the Appendix provides the unit root tests for all variables in the SVAR models considered.[141] Nevertheless, as the price level is differenced twice in this chapter but only once in Chapter 2, it is needed to address this inconsistency. Hence, the baseline model is also estimated without differencing the annual inflation rate. The relevant impulse response

[141] The federal funds rate enters the VAR in levels, even though it is non-stationary for the sample period. This is in line with Bjornland and Leitemo (2009) as the funds rate is also non-stationary for the sample period in their study; however, it is included in levels. Some other studies also include the policy rate in levels, while other variables are differenced to stationarity. For instance, see Bouakez, Essid, and Normandin (2013) or Bjornland and Jacobsen (2010; 2013). Also, the results using the differenced federal funds rate in the baseline model are reported in Figure A3.1 in the Appendix and discussed in Section 3.8.2.

funtions are shown in Figure A3.2 in the Appendix and are briefly discussed in Section 3.8.2.

Thus, the baseline state vector of endogenous variables can be written as follows:

$$Z_t = \left[gap_t, \Delta\pi_t^a, \pi_t^{comp,a}, \Delta sp_t, i_t \right]' \tag{3.15}$$

Following Sims (1992), it is common in the literature to include commodity prices among endogenous variables due to their potential to provide timely information about future inflationary pressures and the current state of the economy for policymakers (Gordon and Leeper, 1994; Sims and Zha, 2006; Vargas-Silva, 2008; D'Amico and Farka, 2011). The data on commodity prices are available daily in the financial market, thus, this strengthens the motivation to use it as an informational variable used by monetary authorities. In other words, many studies use this variable to mitigate the price puzzle.

In order to fully identify the structural model, twenty five restrictions in total must be imposed. The assumption of orthogonal structural shocks and the normalisation of their variance-covariance matrix, i.e. the structural shocks have a unit variance, provide ten and five restrictions, respectively. Thus, additional ten restrictions are required to completely identify the system. In order to recover the structural parameters, the combination of short-run and long-run zero restrictions is used. The short-run zero restrictions can be summarised as follows:

$$\overbrace{\begin{bmatrix} gap_t \\ \Delta\pi_t^a \\ \pi_t^{comp,a} \\ \Delta sp_t \\ i_t \end{bmatrix}}^{Z_t} = D(L) \overbrace{\begin{bmatrix} \beta_{11} & 0 & 0 & 0 & 0 \\ \beta_{21} & \beta_{22} & 0 & 0 & 0 \\ \beta_{31} & \beta_{32} & \beta_{33} & 0 & 0 \\ \beta_{41} & \beta_{42} & \beta_{43} & \beta_{44} & \beta_{45} \\ \beta_{51} & \beta_{52} & \beta_{53} & \beta_{54} & \beta_{55} \end{bmatrix}}^{B_0^{-1}} \overbrace{\begin{bmatrix} \varepsilon_t^{gap} \\ \varepsilon_t^{\pi} \\ \varepsilon_t^{cp} \\ \varepsilon_t^{sp} \\ \varepsilon_t^{mp} \end{bmatrix}}^{\varepsilon_t} \tag{3.16}$$

In line with the standard SVAR literature, it is assumed that output, inflation and commodity prices do not react instantaneously to a monetary policy shock, whilst the monetary policy instrument is allowed to respond contemporaneously to all three variables (Christiano, Eichenbaum and Evans, 2005). Consequently, these macroeconomic variables are ordered above the monetary policy rate in the state vector Z_t. Within the macroeconomic block, recursive causal relationships are assumed. Commodity price inflation is the most responsive of the three variables and reacts contemporaneously to both output gap and consumer price inflation.[142] On the other hand, output gap does not react to any variable contemporaneously. Finally, all variables in the system are allowed to have an immediate impact on real stock returns and the monetary policy rate. However, as in the case of monetary policy, stock returns only affect macroeconomic variables with a lag. Most importantly, two bottom rows of B_0^{-1} imply that the short-run contemporaneous relationship between real stock prices and monetary policy remains unrestricted, i.e. neither β_{45} nor β_{54} is set to zero. These nine short-run zero restrictions are denoted by zeros in the first three rows of the matrix B_0^{-1} in Equation (3.16).

The final (tenth) restriction is imposed on the long-run relationship between the monetary policy rate and real stock prices. It implies that a monetary policy shock has no long-run effect on real stock prices. This restriction is reflected in the long-run response matrix $C(1)$ by setting the infinite sum of relevant lag coefficients in Equation (3.8), $C_{45}(1) = \sum_{k=0}^{\infty} C_{45}(k)$, equal to zero. From Equation (3.13) and the long-run restriction $C_{45}(1) = 0$ it follows that:

$$\begin{bmatrix} C_{11}(1) & C_{12}(1) & C_{13}(1) & C_{14}(1) & C_{15}(1) \\ C_{21}(1) & C_{22}(1) & C_{23}(1) & C_{24}(1) & C_{25}(1) \\ C_{31}(1) & C_{32}(1) & C_{33}(1) & C_{34}(1) & C_{34}(1) \\ C_{41}(1) & C_{42}(1) & C_{43}(1) & C_{44}(1) & 0 \\ C_{51}(1) & C_{52}(1) & C_{53}(1) & C_{54}(1) & C_{55}(1) \end{bmatrix} = \begin{bmatrix} D_{11}(1) & D_{12}(1) & D_{13}(1) & D_{14}(1) & D_{15}(1) \\ D_{21}(1) & D_{22}(1) & D_{23}(1) & D_{24}(1) & D_{25}(1) \\ D_{31}(1) & D_{32}(1) & D_{33}(1) & D_{34}(1) & D_{34}(1) \\ D_{41}(1) & D_{42}(1) & D_{43}(1) & D_{44}(1) & D_{45}(1) \\ D_{51}(1) & D_{52}(1) & D_{53}(1) & D_{54}(1) & D_{55}(1) \end{bmatrix} B_0^{-1}$$
(3.17)

$$D_{41}(1)\beta_{15} + D_{42}(1)\beta_{25} + D_{43}(1)\beta_{35} + D_{44}(1)\beta_{45} + D_{45}(1)\beta_{55} = 0 \quad (3.18)$$

Note that, given the short-run restrictions in Equation (3.16), it then shrinks to:

$$D_{44}(1)\beta_{45} + D_{45}(1)\beta_{55} = 0 \quad (3.19)$$

3.6.4 Augmented model: specification and identification

In the next stage, the baseline model is augmented in line with the recommendations by Brissimis and Magginas (2006). Firstly, the annual commodity price

[142] Such recursive ordering of two macroeconomic variables and commodity prices is in line with Gordon and Leeper (1994) and D'Amico and Farka (2011). The results reported for the baseline model in Section 3.8.2 do not change if the commodity price inflation is placed as the first variable, i.e. it precedes output gap and inflation. In this case, it is assumed that commodity price inflation is contemporaneously exogenous – no variable in the system has an immediate effect on it.

inflation is removed from the state vector.[143] Instead, in order to mitigate the price puzzle, two other variables are included. The composite leading economic indicator published by the Conference Board is included as a measure of future economic activity. In addition, the current-month expected level of the monetary policy rate enters the VAR system as an exogenous variable. The current expected policy rate is measured by the rate on the 1-month federal funds futures contract on the last business day in the previous month (fff_{t-1}). This is a near-perfect proxy for market expectations due to much greater transparency and the openness of monetary policy conduct in the sample period considered here. The augmented model is a VAR-X model.

The additional variables are transformed in line with the baseline specification. The augmented endogenous state vector contains the following variables: the first difference (lagged) in annual change in the composite leading economic indicator ($\Delta lead_{t-1}^{a}$), the output gap (gap_t), the first difference in the annual consumer price inflation ($\Delta\pi_t^a$), the monthly real stock returns (Δsp_t) and the monetary policy rate (i_t):[144]

$$Z_t = \left[\Delta lead_{t-1}^{a}, gap_t, \Delta\pi_t^a, \Delta sp_t, i_t \right]' \tag{3.20}$$

Similarly to the consumer price inflation, the annual change in the leading indicator enters the SVAR in first differences to ensure the stationarity of this variable.[145] While it may be questioned why annual and not monthly changes are used, the robustness analysis is conducted with respect to such data transformation in Section 3.10.7. Finally, to address the issue of inconsistency between Chapter 2 and Chapter 3 in terms of differencing the price level, the main analysis is also carried out with undifferenced annual inflation variable in Section 3.10.8.

[143] As it will be seen later in Section 3.8.2, the inclusion of the commodity price inflation does not eliminate the price puzzle in the baseline model that is quite pronounced. The specification and identification of this model implies that commodity prices could be predicted by the US output gap and inflation. Nevertheless, the estimation results (not reported) indicate that the lags of changes in inflation and the lags of output gap are generally insignificant in the equation for commodity price inflation. According to the p-values of the Granger Causality test, the lags of these two variables do not Granger cause commodity price inflation.

[144] Following Brissimis and Magginas (2006), the leading economic index is included with a lag of one month since some of the data used to compose the Conference Board Leading Economic index for the US is not available until after fifteen days since the end of the month under consideration.

[145] Note that two interest rates in the augmented specification are in levels. This is again consistent with the baseline model and the approach in Bjornland and Leitemo (2009). Also, while they are non-stationary, it is very likely the two series, i.e. the policy rate and expected policy rate, are cointegrated in the longer run. Thus, this should not be a problem.

With respect to the identification of the augmented model, the leading economic indicator is ordered the first as it is not contemporaneously affected by any other variable in the system, while the same long-run restriction applies:

$$\overbrace{\begin{bmatrix} \Delta lead_{t-1}^{a} \\ gap_t \\ \Delta\pi_t^a \\ \Delta sp_t \\ i_t \end{bmatrix}}^{Z_t} = D(L) \overbrace{\begin{bmatrix} \beta_{11} & 0 & 0 & 0 & 0 \\ \beta_{21} & \beta_{22} & 0 & 0 & 0 \\ \beta_{31} & \beta_{32} & \beta_{33} & 0 & 0 \\ \beta_{41} & \beta_{42} & \beta_{43} & \beta_{44} & \beta_{45} \\ \beta_{51} & \beta_{52} & \beta_{53} & \beta_{54} & \beta_{55} \end{bmatrix}}^{B_0^{-1}} \overbrace{\begin{bmatrix} \varepsilon_t^{lead} \\ \varepsilon_t^{gap} \\ \varepsilon_t^{\pi} \\ \varepsilon_t^{sp} \\ \varepsilon_t^{mp} \end{bmatrix}}^{\varepsilon_t} \tag{3.21}$$

3.6.5 Some caveats

In addition to the omitted variables problem, there are several other points to consider when using a SVAR model for monetary policy analysis. Firstly, the form of the monetary policy reaction function and the structure of economy are likely to change over time implying that a constant parameter SVAR may not be suitable, especially for long sample periods (Rudebusch, 1998; Stock and Watson, 2001). Secondly, one must choose an adequate proxy for monetary policy stance (Bernanke and Mihov, 1998). The third point regards the usage of the revised (final) data that is typically used to estimate SVAR models. Hence, the VAR system contains too much information as the final data was not available to monetary policymakers at the time of decision making (Rudebusch, 1998). Finally, structural VAR models use many lags implying the backward-looking monetary policy behaviour (Rudebusch, 1998).

The empirical analysis in this paper is based on the constant-parameter standard SVAR model due to several reasons. Firstly, the samples considered here are the part of the Great Moderation era, known for financial and economic stability. The sample period of main interest starts at the time of major changes in the communication of the Fed’s policy decisions to the public and ends prior to the global financial crisis. Thus, it is reasonable to assume that the Fed’s policy rule and the structure of the US economy have not changed considerably over this period. Secondly, as shown in several studies that use Markov-switching VAR models, the period starting around the mid-1980s is typically described by a single regime (Chang, Chen and Leung, 2011, Lutkepohl and Netsunajev, 2015).[146] With respect to a monetary policy instrument, there is little disagreement in the literature that the federal funds rate is a suitable proxy for the sample periods in this chapter. Finally, the use of the revised data may not pose a severe problem since the measure of future economic conditions, i.e. the composite leading economic indicator, is included in the model with a lag. It acts as a proxy for the information set that policymakers use to deduce future developments in inflation and output.

3.7 Data and sample period

3.7.1 Sample period

The empirical analysis in this chapter is based on monthly US macroeconomic and financial data covering the period from 1985:1 to 2008:12.

The Fed has started to normalise its monetary policy by raising the policy rate target in December 2015 for the first time in almost ten years. Given this, the empirical analysis is focused on conventional monetary policy over the past two decades. The findings in this chapter may prove useful to monetary policymakers as further increases in the federal funds rate target are looming.

The sample period of main interest is 1994:2 – 2007:7.[147]As Fawley and Neely (2014) note, there have been major changes in the Fed's communication of its policy decisions to the public since February 1994. This has led to much greater transparency and predictability of monetary policy conduct that may have had an impact on how financial markets respond to monetary policy shocks. Furthermore, the post-93 period is not commonly used in the standard SVAR literature for the analysis of monetary policy effects. The events around the global financial crisis may have introduced a structural break in the Fed's policy reaction function, hence, the crisis period is excluded from the main analysis.[148]

The empirical analysis begins with a longer sample period. More specifically, the baseline model is firstly estimated over the period 1985:6 – 2007:7, i.e. the Great Moderation era that is known for low and stable inflation and sustained economic growth

[146] Given the sample period in this chapter, it may not be appropriate and/or feasible to employ the heteroscedasticity-based VAR approach with low-frequency data as in Bouakez, Essid and Normandin (2013) or as in Lutkepohl and Netsunajev (2015). This generally requires a very long sample period.
[147] In all cases, the sample period as reported excludes the months reserved for the lags of endogenous variables.
[148] The sensitivity of the results with respect to the inclusion of the period until the zero lower bound is tested in the robustness analysis.

accompanied by a simple and predictable rule underlying the conduct of monetary policy. Equivalently, the augmented model is initially estimated for the period 1989:2 – 2007:7. The sample starts slightly later due to the availability of the federal funds futures market data.

Note that the estimations over the full sample only consider real stock market returns. Meanwhile, the interaction between monetary policy and real returns at the stock market and stock portfolio levels are analysed with respect to the period of interest in this chapter, i.e. 1994:2 – 2007:7.

3.7.2 Macroeconomic variables

The macroeconomic time series used for the empirical analysis include the output gap, the annual consumer price inflation, the annual commodity price inflation, the annual growth in the leading economic indicator, and the federal funds rate. The potential output is constructed by applying the Hodrick-Prescott filter to the series of the log Industrial Production Index (ip_t). Accordingly, the output gap (gap_t) is measured as the deviation of actual production from its potential trend. The annual consumer price inflation (π_t^a) is calculated as an annual change in the log Consumer Price Index for all items (cpi_t). Similarly, the annual commodity price inflation ($\pi_t^{comp,a}$) is computed as an annual change in the log Commodity Research Bureau BLS Spot Index ($comp_t$).

The Conference Board Leading Economic Index for the US (CBLEI) serves as a proxy for the future expected path of economic activity. The annual growth in the leading economic indicator ($lead_t^a$) is calculated as a change in the log CBLEI (lei_t) from a year ago. Finally, the effective federal funds rate (i_t) is used as a proxy for the monetary policy rate. Monthly averages of daily rates are collected from FRED database maintained by the Federal Reserve Bank of St. Louis. Seasonally adjusted data on the IPI and CPI is also provided by FRED, while monthly averages of the commodity price index is constructed using daily data obtained from Datastream database. Monthly CBLEI series for the US are also available in Datastream.

3.7.3 Financial variables

Real stock market returns (Δsp_t) are calculated as a monthly change in the log S&P500 stock market index deflated by the log CPI (sp_t). Monthly averages of the S&P500 index are calculated using daily figures from Datastream.[149] In addition to the aggregate stock market level, the empirical analysis is extended to portfolio-level stock returns. Monthly returns on value-weighted stock portfolios (excluding dividends) are provided by Kenneth R. French.[150] Monthly CPI inflation is subtracted to calculate real stock returns. The data is collected for ten size portfolios formed on the firm's market value (S), ten value portfolios formed on the ratio of book equity value to market equity value (BE/ME), and twenty five double-sorted size-value portfolios. The first decile denotes the size (value) portfolio of the smallest (most growth) stocks, while the tenth decile refers to the size (value) portfolio of the largest (most value) stocks. With respect to the double-sorts, the lowest size (value) quintile represents the smallest (most growth) firms and the highest quintile represents the largest (most value) firms. Finally, the end of month rates on the 1-month federal funds futures contract (fff_t) are obtained from Bloomberg database.

There are several reasons why stock portfolios sorted by the firm's size and book-to-market value ratio are chosen. Firstly, this chapter examines the credit channel of monetary policy transmission that provides two mechanisms how policy decisions could have stronger effects on relatively financially constrained firms as compared to unconstrained firms. As Bernanke and Gertler (1995) explain, the direct monetary policy impact on market interest rates is enhanced through endogenous changes in the external finance premium. Following monetary policy tightening, the external finance premium may increase due to the reduced supply of loans to firms, i.e. the bank lending channel is at work. Alternatively, external funds may become more expensive due to the worsening financial position of borrowers reflected via the changes in their net worth, cash flows, value of collateral and other indicators of balance sheet strength, i.e. the balance sheet channel is active.

The size of a firm is often used as a proxy for financial contraints since small firms are typically more dependent on bank lending and pay higher external finance premium than large firms (Gertler and Gilchrist, 1994; Ehrmann and Fratzscher, 2004; Guo, 2004; Basistha and Kurov, 2008; Kontonikas and Kostakis, 2013; Maio, 2014). Thus, it is

[149] The analysis is based on monthly real stock returns as opposed to monthly observations of annual returns. This is a standard approach in this type of empirical work, especially if stock portfolio data is used.
[150] The data library is accessible via this link: http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html#Research.

sensible to use size-sorted stock portfolios to identify the bank lending mechanism of the credit channel. Similarly, value firms are considered to be relatively more financially contrained than growth firms because they are likely to be less profitable and more heavily reliant on cash flows, thus, more exposed to interest rate changes (Guo, 2004; Kontonikas and Kostakis, 2013; Maio, 2014). Therefore, the stock portfolios formed on the book-to-market value ratio provide an insight into the balance sheet channel of monetary policy transmission.[151] While there may be other proxies for financial constraints, the selected portfolios are in line with the existing similar studies, making it easier to compare the findings.

Secondly, the strand of literature examining differential monetary policy effects on stock prices is closely related to the stock return anomalies analysed in the cross-sectional asset pricing literature. Among most common are the size, value and momentum anomalies. The size anomaly refers to higher average returns on small firms' stocks as compared to large firms (Banz, 1981). It has also been shown that value stocks tend to earn higher average returns, i.e. the value premium (Basu, 1983; Fama and French, 1992; Lakonishok, Shleifer and Vishny, 1994). The momentum premium implies that winner stocks that earned higher returns in the recent past continue to deliver higher returns in the future (Jegadeesh and Titman, 1993). Fama and French (1992; 1993) demonstrate that the cross-section of US excess stock returns can be explained by the excess market return and two common risk factors related to the firm's size and the BE/ME ratio that are omitted in the standard capital asset pricing model (CAPM).[152] Interestingly, the beta associated with the market risk is similar across stock portfolios in the three-factor model; however, the slopes on the two factors related to the size and book-to-market value vary much more.[153] They argue that higher average returns must reflect a compensation for a higher risk associated with smaller and value stocks.

Subsequently, a great number of studies have attempted to provide the economic interpretation of these risk factors.[154] Several studies indicate the potential of a monetary

[151] Alternative proxies for value stocks, that can also be used to analyse this channel, are cash-flow-to-price, earnings-to-price, and dividend-to-price ratios.

[152] Fama and French (2015, 2016) add two more risk factors in the model for stock returns to capture investment and profitability risk premiums.

[153] On the other hand, Campbell and Vuolteenaho (2004) argue that the size and value stock return anomalies could be explained by the two-beta model where a single beta in the CAPM is divided into the component reflecting cash flow news (bad beta) and the component related to discount rate news (good beta). They find that value and small stocks are associated with higher cash-flow betas as compared to growth and large stocks. Thus, this may explain higher average returns on value and small stocks.

[154] Fama and French (1995) argue that firms with a low book-to-market value ratio earn sustained profits, while firms with a high ratio tend to be relatively financially distressed due to low profitability. Others find that the size and value factors positively predict future real gross domestic product (Liew and Vassalou, 2000; Vassalou, 2003) or, alternatively, they represent the default risk (Vassalou and Xing, 2004).

policy risk factor to explain stock market anomalies since many macroeconomic variables, that may be related to Fama-French risk factors, are also related to monetary policy conditions (Thorbecke, 1997; Jensen and Mercer, 2002; Hahn, O'Neill and Reyes, 2004; Arshanapalli, Fabozzi and Nelson, 2006; Kontonikas and Kostakis, 2013; Lioui and Maio, 2014; Maio, 2014). Hence, this chapter also investigates whether the Fed's policy can shed some light on the two commonly analysed size and value anomalies in stock returns.[155]

3.8 Empirical modelling and some initial results

As recommended by Sims and Zha (1999), throughout this chapter the impulse response functions are reported together with the probability bands represented as 0.16 and 0.84 fractiles, i.e. 68% probability bands.[156] The median value is chosen as the central measure of impulse responses. The Monte Carlo integration with 10,000 draws is used to obtain the Bayesian simulated distribution of impulse response functions using the approach for just-identified VAR systems. The draws are made from the posterior distribution for VAR parameters and residuals under the standard uninformative ("flat") prior for a multivariate regression model (Doan, 2015).

3.8.1 Replication of Bjornland and Leitemo (2009)

Prior to estimating the baseline SVAR model, this chapter attempts to replicate the findings in Bjornland and Leitemo (2009).[157] The original sample period spans 1984:6 – 2002:12 and four lags of each endogenous variable are included. Figures 3.1-3.2 depict the replication of the impulse response functions (IRFs) in Figure 2 presented in Bjornland and

Furthermore, the value factor is shown to be positively related to the term spread and the size factor negatively relates to the default spread (Hahn and Lee, 2006; Petkova, 2006). In contrast, Aretz, Bartram and Pope (2010) find a negative relationship between the value factor and expected output growth, whilst the size factor has a positive loading on survival probability and on the average level of the term structure.

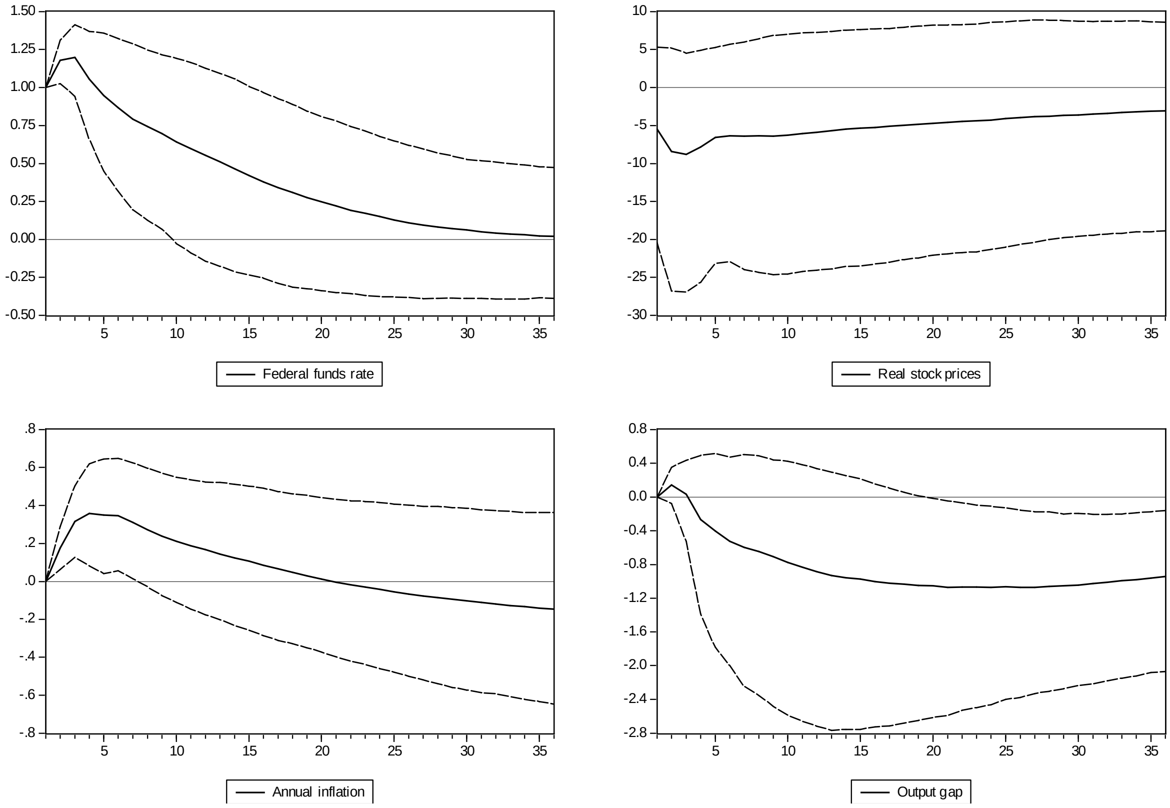
[155] It may also be of interest to investigate the stock portfolios sorted by the "good" and "bad" market beta. As highlighted in Campbell and Vuolteenaho (2004), the two-beta asset pricing model can explain the size and value anomalies in stock returns. However, such portfolios are not readily available, thus, this analysis is out of the scope of this chapter.

[156] According to Sims and Zha (1999), the traditional symmetric confidence intervals, reported as one or two standard errors around the point estimate of an impulse response function, may be misleading. Such error bands confound information regarding the location of coefficient values with information about overall model fit. They show that the Bayesian posterior probability bands, simulated using Monte Carlo integration, may be more useful than the confidence intervals based upon the estimates of standard errors. The fractiles correspond to a one standard deviation if the standard error bands were used.

[157] The original dataset and data transformation are used. The dataset and the RATS code for the replication are provided by Tom Doan and are available at https://estima.com. I have amended the original version of the RATS code accordingly to produce the results reported in this chapter.

Leitemo (2009).[158] Figure 3.1 demonstrates the responses of the federal funds rate, real stock prices, annual inflation and output gap over the 36-month horizon after a positive 1-percentage-point shock in the funds rate. Figure 3.2 shows the responses of the same variables to a 1 % increase in real stock prices (market level). While the shape of impulse response functions is similar to those in the original study, the probability bands reported here are much wider. For instance, the decline in real stock prices following a contractionary monetary policy shock is never statistically significant. Furthermore, the response of the funds rate to a stock price shock only becomes statistically significant with a lag of around five months as opposed to an immediate significant response as reported in the original study.

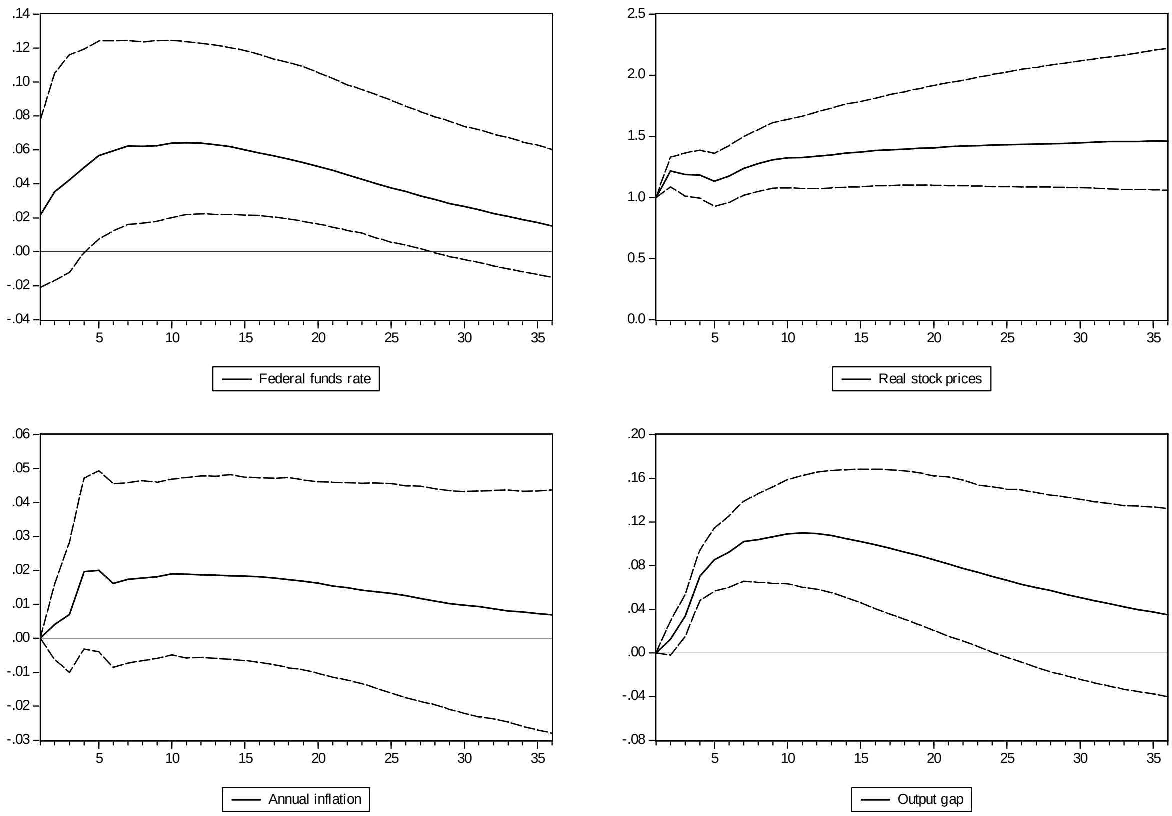
Figure 3.1: Impulse responses to FFR shock - replication of BL2009



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation and output gap following a 1-percentage-point contractionary monetary policy shock. The SVAR model is specified as in Bjornland and Leitemo (2009) and is estimated over the sample period 1984:6 – 2002:12 including 4 lags. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 2500 draws.

[158] The Bayesian simulated distribution is obtained using Monte Carlo integration with 2500 draws as in the original paper. The median value is chosen as the central measure of impulse responses.

Figure 3.2: Impulse responses to SP shock - replication of BL2009



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation and output gap following a positive 1% stock price shock. See also Figure 3.1 notes.

It is important to note that the structural model depends on the new set of parameters for each Monte Carlo draw and one has to make sure that model's parameters are reset accordingly prior to the generation of new impulse response functions. Nevertheless, some empirical work in the literature erroneously identifies the structural model using the original OLS estimates of VAR coefficients in each draw (Doan, 2015). This may explain why it is not possible to replicate the results in Bjornland and Leitemo (2009). Indeed, if the original OLS parameters are used in every draw, then their findings can be replicated. Thus, their argument that there is the strong interdependence between the US stock market and monetary policy appears to be weak.

3.8.2 The baseline model

This section estimates the baseline SVAR model as specified in Section 3.6.3 over the sample period 1985:6 – 2007:7. Panel A of Table A3.1 ine the Appendix provides the augmented Dickey-Fuller unit root test results for the variables in the vector Z_t. The model is estimated using six lags of endogenous variables. Also, two dummies are included as exogenous variables that take value of one in October 1987 and September 2001, respectively, to account for the stock market crash and zero otherwise. The model satisfies the stability condition, i.e. all inverse roots lie inside the unit circle, and is invertible. Also, there is no autocorrelation or heteroscedasticity present in the residuals.[159]

The impulse responses of the federal funds rate, real stock prices, annual inflation and output gap to a contractionary monetary policy shock are reported in Figure 3.3. The dashed lines represent 68% probability bands and the solid line denotes the central measure of impulse responses, i.e. the median value. The shock has a positive and persistent impact on the federal funds rate as it remains well above the pre-shock level over the horizon of 36 months. Initially, annual inflation rate continues to increase and this reaction is statistically significant for several months. Inflation peaks at around 0.6% during the ninth period and then starts to decline; however, the response is still positive after three years following the shock. The evident price puzzle indicates a potential misspecification of the baseline model. Following monetary policy tightening, higher cost of borrowing, lower asset prices and reduced wealth should lead to a decline in consumption and investment in the economy. As a result, the price level is expected to decline and inflation rate should be reduced. Furthermore, output gap also responds positively to unexpected monetary policy tightening but it eventually declines below the initial level after around two years. However, the response is generally statistically insignificant. This persistent increase in output is not in line with the transmission mechanism of monetary policy that predicts a decline in aggregate demand. Overall, the findings for inflation and output are not consistent with predictions by standard economic theories.[160]

Following an unexpected increase in the federal funds rate by 1 percentage point, real stock prices decline sharply on the impact. The initial drop of approximately 11.30% is both statistically significant and economically important. A typical surprise change in the target for the federal funds rate of 25 basis points is associated with approximately 2.83% change in real stock prices. The magnitude of the response is somewhat greater than what is typically found in similar studies. For instance, D'Amico and Farka (2011) find that a 25-basis-point unexpected increase in the funds rate leads to 1.25% decline in the stock market returns, while Li, Iscan and Xu (2010) demonstrate that such a policy shock results in real stock prices declining by 0.55%. On the other hand, the pronounced real stock

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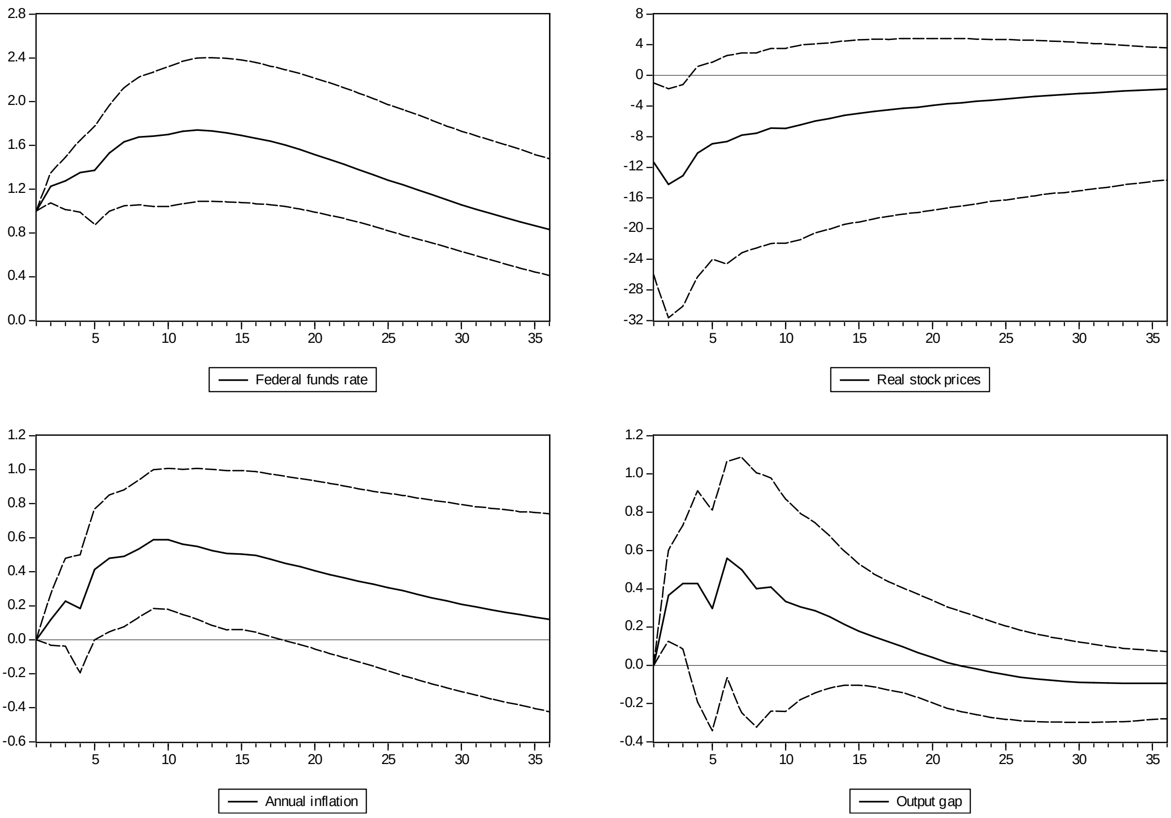
[159] LM test is used to detect serial correlation in the residuals of the estimated model and, the White test is used to identify heteroscedasticity.

[160] As Figure A3.1 in the Appendix shows, these counter-intuitive effects of a contractionary monetary policy shock on macroeconomic variables are even more pronounced if the federal funds rate is differenced in the state vector. Equally, the signs of the model misspecification remain if the annual consumer inflation is not differenced as demonstrated in Figure A3.2 in the Appendix.

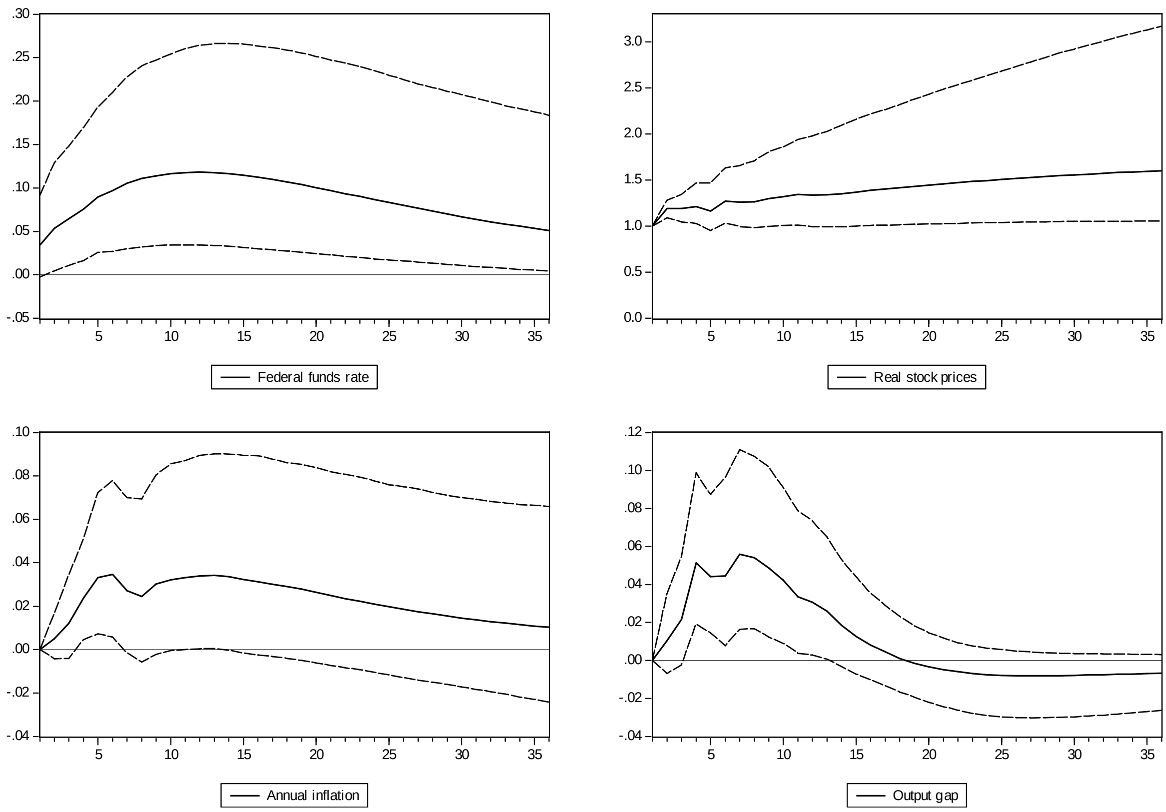
market decline in Figure 3.3 is consistent with the studies using the restriction that monetary policy has no long-run effects on real stock prices in the combination with standard short-run restrictions (Bjornland and Leitemo, 2009; Bjornland and Jacobsen, 2013). The monetary policy effect on real stock prices remains statistically significant for another two months following the shock. The contractionary policy shock can lower real stock prices via its negative impact on expected future cash flows. For instance, higher market interest rates increase interest payments of firms and also have a dampening effect on overall economy leading to lower demand for firms' production. In addition, monetary policy tightening can depress the stock market through a higher discount rate used to discount future expected cash flows. The negative response of the stock market to surprise monetary policy tightening is also reported by many other studies (Thorbecke, 1997; Bernanke and Kuttner, 2005; Bjornland and Leitemo, 2009; D'Amico and Farka, 2011; Bjornland and Jacobsen, 2013).

Figure 3.3: Impulse responses to FFR shock - baseline model



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The baseline SVAR model is estimated over the sample period 1985:6 – 2007:7 including 6 lags. The state vector contains the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the annual commodity price inflation ($\pi_t^{comp,a}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). Two dummy variables that take value of one in 1987:10 and 2001:9, respectively, and zero otherwise are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

Figure 3.4: Impulse responses to SP shock - baseline model



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure 3.3 notes.

The structural stock price shock identified in the model reflects the variation in equity prices that is not explained by current macroeconomic data. Figure 3.4 shows the dynamic impact of a positive 1% shock in real stock prices at market level on the four endogenous variables. With respect to macroeconomic variables, both inflation and output gap increase and the magnitude of this positive effect is moderate. The unexpected increase in real stock prices by 10% would lead to an increase in inflation rate by 0.33 percentage points in five months. Meanwhile, output gap would be about 0.56 percentage points higher after 7 months. The impulse responses turn statistically significant after several periods with output gap response being significant for a longer period. These findings are overall consistent with the view that higher stock prices boost investment through the Tobin's Q and lead to higher consumption spending through changes in wealth subsequently resulting in higher aggregate price and output levels.

As discussed in Chapter 1, monetary policymakers may have an incentive to monitor stock prices due to their forward-looking nature. Asset prices could possibly provide some information about future developments in consumption, investment spending and financial markets that has not yet been reflected in current macroeconomic variables.

Figure 3.4 indicates that the positive stock price shock results in tighter monetary policy stance. The initial reaction of the federal funds rate is positive (three basis points) and statistically insignificant, but it turns significant in the second period. The funds rate rises to eleven basis points within one year and then starts slowly returning to its initial level. Thus, an unexpected 10% rise in real stock prices results in the federal funds rate increasing by approximately 0.3% initially and by 1.1% one year after the shock. The finding of a strong monetary policy response to stock prices is consistent with a forwad-looking inflation-targeting central bank that sets the policy rate in order to curb future inflationary pressures in the economy reflected in rising stock prices. It is possible that the macroeconomic effects of the stock price shock could be larger if monetary policy did not react to stock prices in this manner. On the other hand, unexpectedly higher stock market valuation could be indicative of an asset price bubble that may also prompt policymakers to take action to preserve financial and economic stability. Nevertheless, this model does not allow to distinguish the bubble component of stock price shock. The positive and significant Fed's reaction to stock price developments is in line with findings in the literature (Rigobon and Sack, 2003; Bjornland and Leitemo, 2009; D'Amico and Farka, 2011; Bjornland and Jacobsen, 2013).

In summary, the baseline findings indicate that unexpected monetary policy tightening is associated with a sharp decline in the stock market, whilst the federal funds rate rises following a positive shock to real stock prices. This evidence provides support for the interdependence between monetary policy and the stock market in the US. However, it is likely that the baseline model is not well specified as indicated by the price puzzle and the anomalous response of output gap to monetary policy shocks. This motivates to improve the baseline model in order to generate the impulse response functions that are in line with economic theory predictions. Therefore, the SVAR model is augmented following the suggestions by Brissimis and Magginas (2006).

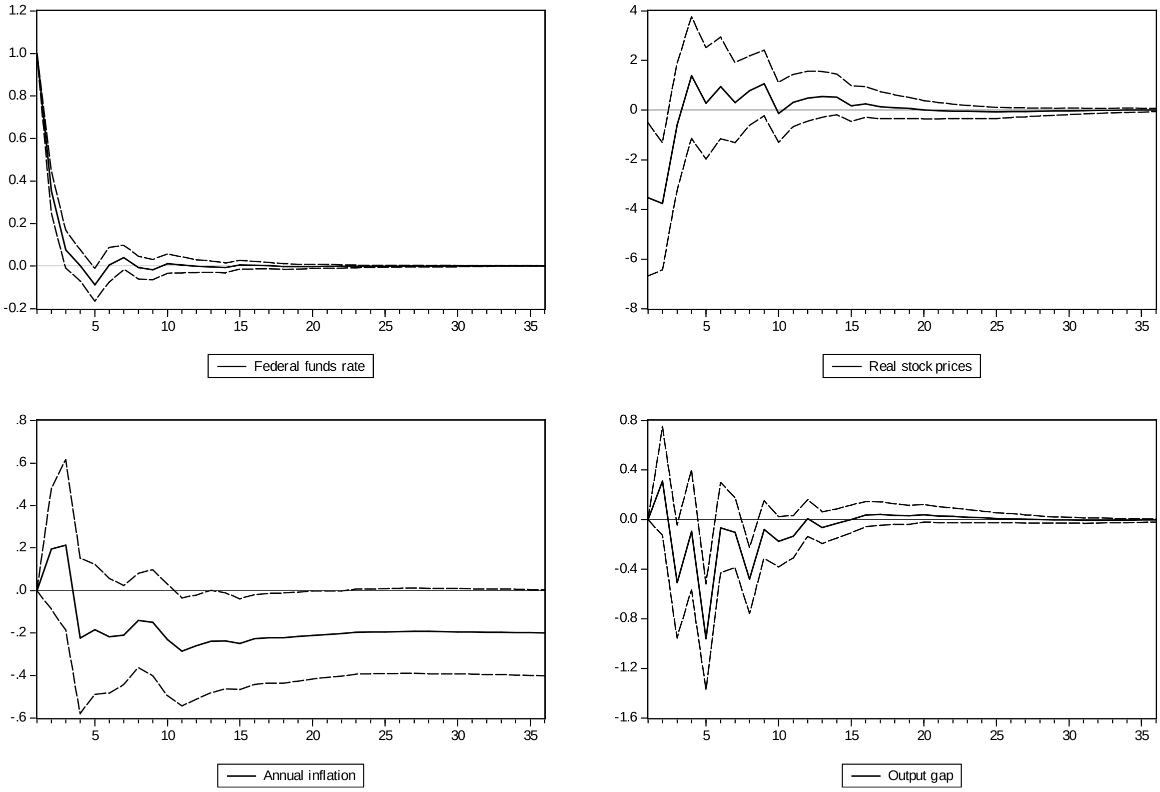
3.8.3 The augmented model: full-sample analysis of stock market returns

The augmented model is initially estimated with real stock market returns over the sample period 1989:2 – 2007:7 using six lags and the dummy variable to account for the stock market crash in September 2001.[161] The model satisfies the stability condition and is

[161] Panel B of Table A3.1 in the Appendix reports the relevant unit root test results. Due to the availability of the federal funds futures market data, the sample starts later as compared to the baseline model.

invertible. Also, there is no evidence of serial autocorrelation or heteroscedasticity in the model's residuals.

Figure 3.5: Impulse responses to FFR shock - augmented model



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1989:2 – 2007:7 including 6 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^a_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

The impulse responses of the federal funds rate, real stock prices, annual inflation and output gap to a contractionary monetary policy shock are shown in Figure 3.5. As compared to the baseline model, the impact of monetary policy is quite different. Firstly, the positive response of the federal funds rate peaks on the impact and then dies out quickly. This is in sharp contrast to the persistent increase in the funds rate depicted in Figure 3.3. Thus, the monetary policy innovation paradox is avoided in contrast to the baseline model (Dueker, 2002).[162] Secondly, the dynamics of macroeconomic variables

[162] This paradox refers to the impulse response function of the federal funds rate to a monetary policy shock, i.e. a shock to itself, that continues upwards for several more periods before starting to decay. This implies subsequent responses of policymakers to a policy surprise in the same direction. Such additional actions

following monetary policy tightening are now in line with standard economic theory. The initial increase in annual inflation over the next several months is short-lived and statistically insignificant. After about a year, annual inflation has declined by approximately 0.29 percentage points indicating a fall in the price level, as it would be expected. The impulse response also turns statistically significant around this time, albeit temporarily. An unexpected increase in the funds rate also depresses real economic activity. After five months, output gap has fallen significantly by around 0.95 percentage points. Nevertheless, this effect is short-lived.

The magnitudes of the responses of macroeconomic variables are economically meaningful and generally in line with VAR-based studies for monetary policy analysis over sample periods since the mid-1980s (Boivin, Kiley and Mishkin, 2010; Bjornland and Jacobsen, 2013). Nevertheless, using a FAVAR-based analysis and the estimation of DSGE model, Boivin, Kiley and Mishkin (2010) demonstrate that monetary policy effects on prices and real activity are weaker in more recent period as compared to the period before the 1980s.

With respect to the stock market reaction, the findings are more consistent with the baseline model. The initial impact of a positive policy shock on real stock prices is negative and statistically significant for the first two periods. The magnitude of the impulse response is smaller as compared to the baseline results but still quantitatively important.[163] Real stock prices drop by 3.53% on the impact and this estimate is now closer to what is typically reported in other studies (Li, Iscan and Xu, 2010; Cheng and Jin, 2013; Kontonikas and Kostakis, 2013; Maio, 2014). The stock market almost entirely recovers in the third period following the shock.

Figure 3.6 shows the dynamic effects of a 1% positive shock to real stock prices (at aggregate level). As previously, the findings differ from the baseline results. There is no clear and statistically significant impact on annual inflation. On the other hand, the response of output gap is positive and turns statistically significant for several periods before completely dying out after about one year. Surprisingly, the initial reaction of the federal funds rate to an unexpected increase in real stock prices is negative, although it is statistically insignificant. In order to act pre-emptively in response to expected future increase in aggregate demand, a central bank should raise its policy rate not reduce it. Eventually, the funds rate increases and peaks in the fifth period. With a delay of several

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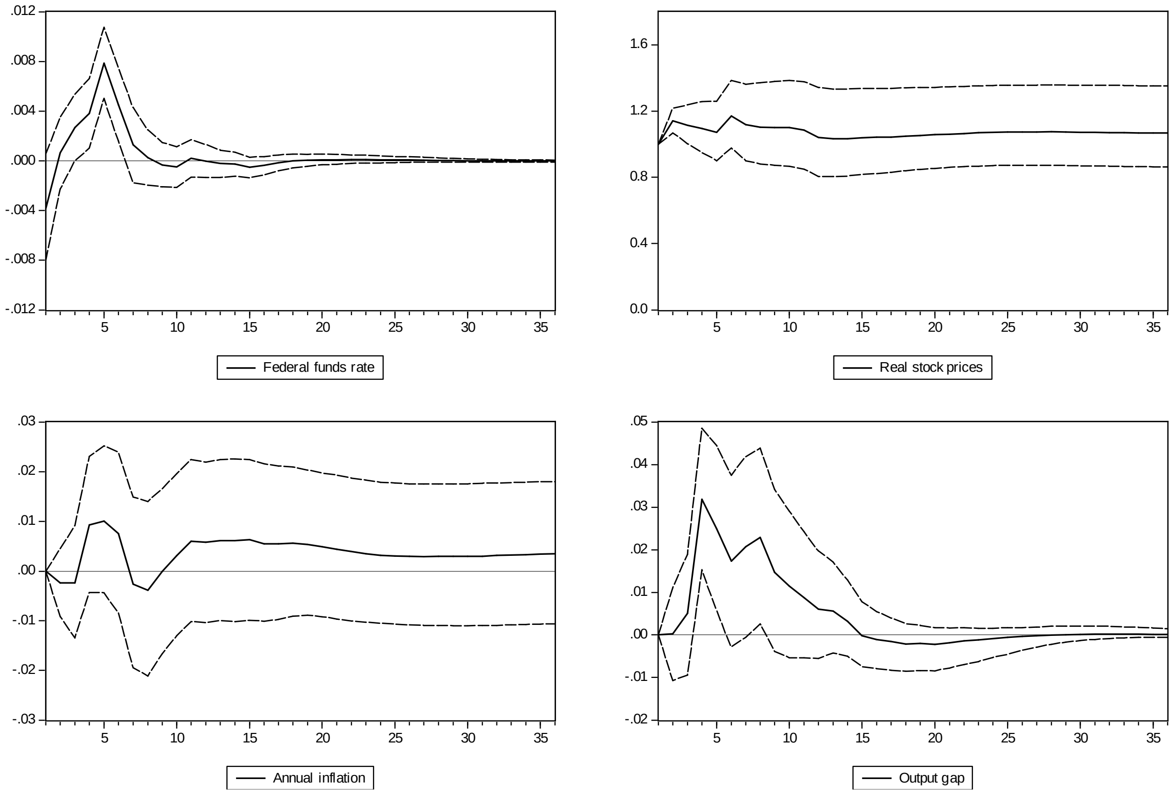
would not have a surprise element and they would not reflect the response to economic developments (see Dueker, 2002).

[163] If the baseline model is also estimated over the same sample, i.e. 1989:2 – 2007:7, it can be noted that the change in the size of the stock market response is driven by the change of the sample period. However, even then the baseline model generates the counter-intuitive monetary policy effects on macroeconomic variables.

months, the reaction also becomes statistically significant. This slow policy response to unexpectedly higher stock prices could possibly indicate that policymakers wait for other signs of economy heating up before acting solely on the information from the stock market. Nevertheless, it is important to note that the absolute size of the response is small and much smaller than in the baseline case. At the peak, the funds rate is only approximately 0.8 of a basis point above the initial level. Given a 10% surprise increase in real stock prices, this translates into the federal funds rate going up by eight basis points after five months To compare, D'Amico and Farka (2011) show that a 1% (10%) increase in stock market returns triggers a policy rate response of 2.6 (26) basis points on average over the period 1983 - 2006. Meanwhile, Bjornland and Jacobsen (2013) estimate that the federal funds rate increases by about 50 basis points following a positive 10% shock to real stock prices. Thus, the evidence of a strong Fed's response to the stock market is somewhat weak in the analysis here.

Figure 3.6: Impulse responses to SP shock - augmented model



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure 3.5 notes.

Overall, the augmented model produces sensible results with respect to monetary policy effects on traditional macroeconomic variables. This implies that the model is well specified or at least better specified than the baseline model. The results also confirm the negative impact of contractionary monetary policy on the stock market. Nevertheless, the findings do not provide convincing evidence for the strong Fed's reaction to the US stock market on a monthly basis.[164] This is in line with the argument by Bernanke and Gertler (1999) that monetary policy should focus on macro-fundamentals, rather than stock market developments, except insofar as they predict changes in relevant economic variables. After including two forward-looking variables in the SVAR model, it appears that real stock prices do not contain much relevant information to policymakers in addition to macroeconomic data and expectations about the future paths of monetary policy and economic activity. These results are in contrast to other studies that find the evidence for the strong interdependence between the US stock market and monetary policy (Rigobon and Sack, 2003; Bjornland and Leitemo, 2009; D'Amico and Farka, 2011; Bjornland and Jacobsen, 2013). On the other hand, Bouakez, Essid, and Normandin (2013) also find little evidence that the stock market has a direct influence over the Fed's monetary policy in 1982 – 2007.

3.9 Empirical results

This section presents the main empirical analysis that is based on the estimation of the augmented model over the period 1994:2 – 2007:7. In addition to real stock market returns, this section also considers real returns on stock portfolios formed on the firm's size (market value) and BE/ME ratio, i.e. size- and value-sorted portfolios. The augmented model is estimated with four lags of each endogenous variable and also includes the dummy variable to account for the stock market crash in September 2001. As previously, the model is found to satisfy the stability condition and, hence, is invertible. There is no serial autocorrelation or heteroscedasticity present in the residuals. Initially, the model is estimated with stock market returns and, subsequently, it is estimated replacing stock market returns with stock returns on each portfolio in turn.

Before turning to the discussion of the results, it is worthwhile having a look at the summary statistics for monthly real stock returns at market and portfolio levels provided in Table 3.1. Monthly real stock market return is on average 0.51% per month over the

[164] Note that this result is not directly comparable to the findings in Chapter 1. Firstly, annual stock price inflation is considered in Chapter 1 as opposed to monthly stock returns in the current chapter. Secondly, the policy rule in the SVAR is not exactly the same as the augmented Taylor rules estimated in the first empirical part of the thesis. On the other hand, the findings in Chapter 3 are not overall inconsistent with the results in Chapter 1 as it has been demonstrated that the Fed's response to financial indicators is largely driven by the period of the global financial crisis that is excluded from the analysis in this chapter.

sample period. This figure lies within the range of stock returns across size and value decile portfolios. The portfolio of the smallest stocks (S1) earns higher average return of 0.87% as compared to 0.44% return on the portfolio of the largest stocks (S10). Nevertheless, the decline in returns is not monotonic across the remaining deciles. While the most value firms (BM10) have 0.71% monthly average return, the return for the most growth firms (BM1) is only 0.46% per month. Again, the increase in average returns across the remaining deciles is not monotonic. In general, returns on smaller stocks are also more volatile than returns on larger stocks. The dispersion is somewhat smaller across value-sorted portfolios in terms of the standard deviations of returns. With respect to double-sorted size-value portfolios, the smallest stocks have higher average returns than the largest stocks within all but the most growth quintile. Within three lowest size quintiles, the most value portfolios (BM5) have higher average returns as compared to the most growth portfolios (BM1). Overall, Table 3.1 provides some evidence for the size and value premiums in average stock returns.

3.9.1 Stock market

To begin with, it may be useful to look at the estimated parameters of the augmented VAR with stock market returns reported in Table 3.2. This provides some insight into how good the model is and what role two forward-looking variables play.[165] In the reduced form model, every endogenous variable is defined as the linear combination of its own lags, lags of other endogenous variables and any exogenous variables. It does not reveal any structural relationships between the variables and it is generally very difficult to interpret the estimated parameters, especially if the VAR contains many variables and lags.

[165] The main analysis is focused on the period since 1994. Thus, the respective estimates for this model over the sample period 1989:2 – 2007:7 are not provided in Section 3.8.3. Nevertheless, the overall conclusions are similar if the sample starts in early 1989.

Table 3.1: Summary statistics for real stock market and stock portfolio returns

Returns	Mean	Median	Max	Min	Std. Dev.	Skewness	Kurtosis
Panel A: Market returns							
Market	0.5051	0.5769	10.3500	-11.5131	3.3121	-0.5799	4.7038
Panel B: Size portfolio returns							
S1	0.8716	1.2389	25.4848	-23.5809	6.0805	-0.1763	5.9017
S2	0.7781	0.8644	22.5794	-26.0384	6.3909	-0.3201	5.2604
S3	0.6742	1.2575	16.5870	-23.8594	5.8442	-0.7115	4.7287
S4	0.5297	1.2584	15.2194	-21.9008	5.5935	-0.6852	4.5987
S5	0.6105	1.2885	13.4505	-22.8125	5.4484	-0.7197	4.6224
S6	0.5344	1.1608	9.2761	-22.4868	4.9117	-0.9349	5.2336
S7	0.7682	1.1055	12.0566	-20.9726	4.7182	-0.7945	5.0588
S8	0.5982	0.9888	11.8445	-19.0088	4.8112	-0.6720	4.7842
S9	0.6661	1.3482	8.6728	-16.3273	4.1423	-0.7620	4.3370
S10	0.4442	0.8267	9.4523	-15.8581	4.2214	-0.6687	4.1071
Panel C: Value portfolio returns							
BM1	0.4577	0.4172	10.2568	-16.3508	4.9469	-0.6225	3.8889
BM2	0.5472	0.9967	10.6551	-15.0421	4.3103	-0.5623	3.8148
BM3	0.5819	0.7698	9.3871	-18.7073	4.2674	-0.8412	5.2820
BM4	0.6833	1.0704	9.4549	-20.8372	4.2739	-0.8870	6.3542
BM5	0.6299	1.1240	11.0589	-20.8495	4.1750	-0.8929	6.6704
BM6	0.5843	0.9973	9.4532	-18.9376	4.0665	-0.9318	5.7835
BM7	0.7335	1.2867	9.1641	-15.1118	3.9496	-0.7127	4.2423
BM8	0.5383	0.8326	8.3701	-14.9609	3.7645	-0.8318	5.2637
BM9	0.5630	1.2586	11.6249	-14.7639	4.1250	-0.7610	4.5122
BM10	0.7118	1.2144	13.1679	-16.8752	4.4902	-0.7757	5.0097
Panel D: Size-value portfolio returns							
S1BM1	-0.1033	0.6952	33.1274	-31.8679	8.9299	-0.0566	5.0456
S1BM2	0.8172	1.2172	32.1863	-25.0456	7.2117	0.0519	6.4112
S1BM3	0.9703	0.7497	23.5284	-22.3369	5.3760	-0.1290	6.2370
S1BM4	1.1807	1.2297	17.1674	-21.5904	4.8546	-0.4280	6.0337
S1BM5	1.2663	1.1439	12.6727	-19.9188	5.0299	-0.7088	4.8459
S2BM1	0.1911	1.2405	23.8349	-28.9040	7.7012	-0.4791	4.1289
S2BM2	0.5807	0.8981	15.3305	-22.7874	5.5457	-0.7040	4.9433
S2BM3	0.8113	0.9586	11.8596	-19.6145	4.6044	-0.8306	5.4825
S2BM4	0.8517	1.0302	9.9992	-20.4566	4.8346	-1.0513	5.8352
S2BM5	0.7310	1.2364	11.3377	-23.9974	5.2303	-1.0636	6.4194
S3BM1	0.2846	1.5087	21.6374	-27.0543	7.1273	-0.5712	4.3688
S3BM2	0.6232	0.8021	11.9570	-22.9632	4.9662	-0.8215	5.5824
S3BM3	0.7676	1.0435	11.6556	-20.9603	4.3947	-0.9321	6.2321
S3BM4	0.6217	0.9337	13.9293	-17.9993	4.4316	-0.7022	5.6478
S3BM5	0.9283	1.0355	12.0463	-21.5277	4.7605	-1.1490	7.1917
S4BM1	0.5838	1.0261	22.6112	-22.4368	6.3257	-0.3357	4.9270
S4BM2	0.7243	0.7831	10.6660	-23.0261	4.5589	-0.9805	6.8860
S4BM3	0.7572	1.0997	14.2383	-20.1385	4.3867	-0.7833	6.1472
S4BM4	0.7890	1.2554	10.4586	-15.1974	4.3112	-0.8309	5.1143
S4BM5	0.5764	1.1767	11.2998	-19.6824	4.5923	-0.9892	5.9422
S5BM1	0.5099	0.6224	9.6368	-14.4518	4.5199	-0.5356	3.5678
S5BM2	0.6276	0.9408	9.2958	-18.7554	4.1957	-0.7968	5.3493
S5BM3	0.5049	0.9565	9.6211	-19.6996	4.1679	-0.8706	5.6721
S5BM4	0.4946	0.8181	13.1633	-13.4985	4.0664	-0.5422	4.0230
S5BM5	0.4010	1.3450	12.5167	-16.3076	4.7603	-0.7461	4.0835

Notes: This table reports the summary statistics for monthly real stock returns on the stock market (Panel A), ten size-sorted decile portfolios (Panel B), ten value-sorted decile portfolios (Panel C) and twenty five double-sorted size-value quintile portfolios (Panel D) for the sample period 1993:10 – 2007:7.

Table 3.2: The OLS estimation of the reduced-form augmented model

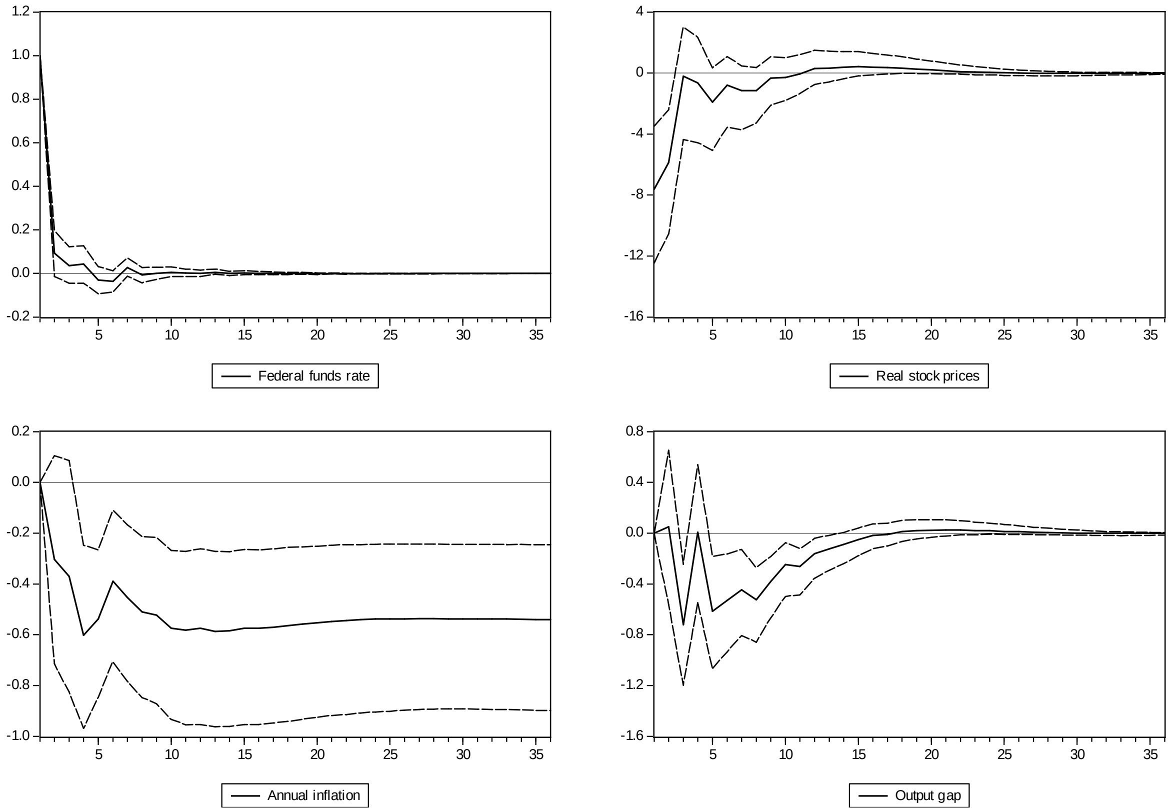
	$\Delta lead^a_{t-1}$	gap_t	$\Delta\pi^a_t$	Δsp_t	i_t
$\Delta lead^a_{t-2}$	0.031 (0.084)	0.109* (0.057)	0.035 (0.038)	-0.459 (0.401)	0.007 (0.010)
$\Delta lead^a_{t-3}$	0.146* (0.084)	-0.039 (0.056)	0.029 (0.038)	0.155 (0.400)	-0.012 (0.010)
$\Delta lead^a_{t-4}$	0.189** (0.081)	0.047 (0.054)	-0.012 (0.037)	-0.422 (0.385)	-0.012 (0.010)
$\Delta lead^a_{t-5}$	0.037 (0.079)	0.013 (0.053)	0.045 (0.036)	-0.030 (0.375)	0.009 (0.009)
gap_{t-1}	0.362*** (0.127)	0.688*** (0.085)	0.027 (0.058)	-1.016* (0.604)	0.011 (0.015)
gap_{t-2}	-0.397** (0.156)	0.196* (0.104)	0.024 (0.071)	1.436* (0.739)	-0.028 (0.019)
gap_{t-3}	-0.068 (0.155)	0.090 (0.104)	-0.059 (0.071)	-0.800 (0.737)	0.025 (0.018)
gap_{t-4}	-0.088 (0.134)	-0.155* (0.090)	0.088 (0.061)	-0.291 (0.637)	-0.008 (0.016)
$\Delta\pi^a_{t-1}$	-0.053 (0.187)	0.090 (0.125)	0.236*** (0.085)	-0.931 (0.886)	-0.027 (0.022)
$\Delta\pi^a_{t-2}$	0.087 (0.193)	0.044 (0.129)	-0.466*** (0.088)	-0.442 (0.918)	0.021 (0.023)
$\Delta\pi^a_{t-3}$	-0.046 (0.195)	0.168 (0.130)	0.000 (0.089)	-0.007 (0.925)	-0.004 (0.023)
$\Delta\pi^a_{t-4}$	-0.042 (0.188)	0.270** (0.126)	-0.073 (0.086)	-0.335 (0.892)	0.001 (0.022)
Δsp_{t-1}	0.068*** (0.018)	0.008 (0.012)	0.001 (0.008)	0.170** (0.083)	0.001 (0.002)
Δsp_{t-2}	0.005 (0.019)	-0.012 (0.013)	-0.003 (0.009)	0.014 (0.091)	0.002 (0.002)
Δsp_{t-3}	0.017 (0.019)	0.029** (0.013)	0.005 (0.009)	-0.045 (0.090)	0.004* (0.002)
Δsp_{t-4}	-0.023 (0.019)	-0.007 (0.013)	0.007 (0.008)	0.027 (0.089)	0.005** (0.002)
i_{t-1}	-1.664* (0.871)	0.120 (0.583)	-0.300 (0.396)	2.863 (4.133)	0.103 (0.104)
i_{t-2}	-0.432 (0.748)	-0.618 (0.501)	0.084 (0.340)	4.273 (3.551)	0.036 (0.089)
i_{t-3}	0.840 (0.744)	0.740 (0.498)	-0.212 (0.338)	-3.373 (3.530)	0.054 (0.088)
i_{t-4}	0.050 (0.450)	-0.443 (0.301)	0.161 (0.205)	-0.730 (2.135)	-0.087 (0.054)
c	0.086 (0.168)	-0.169 (0.113)	0.087 (0.076)	-0.342 (0.798)	0.038* (0.020)
fff_{t-1}	1.164** (0.545)	0.242 (0.365)	0.245 (0.248)	-2.832 (2.585)	0.880*** (0.065)
$D^{2001:9}$	0.634 (0.724)	-0.390 (0.485)	-0.194 (0.330)	-10.739*** (3.437)	-0.391*** (0.086)
Adj. R^2	0.385	0.829	0.160	0.092	0.998
F-statistic	5.573	36.358	2.392	1.739	3481.771
P-value	0.000	0.000	0.001	0.029	0.000

Notes: This table reports the OLS estimates of the reduced-form VAR together with standard errors in parentheses over the sample period 1994:2 – 2007:7 including 4 lags. The endogenous state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^a_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the monthly real stock returns (Δsp_t) and the federal funds rate (i_t). The dummy variable ($D^{2001:9}$) that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate as of last business day of previous month (fff_{t-1}) are included as exogenous variables. The last two rows report the *F*-statistics to test the null hypothesis that all of the regression parameters are zero and the corresponding *p*-values ***, **, * denote 1%, 5% and 10% level of significance, respectively.

As it is shown in Table 3.2, the majority of the coefficients are statistically insignificant and different lags of the same endogenous variable tend to have opposing signs within the same equation. With respect to the leading economic indicator, some of its own lags and some lags of output gap and stock returns are statistically significant. The expected level of the interest rate is also significant. The output gap can be predicted by some lagged values of its own, leading indicator, inflation and stock returns. With respect to inflation, only some lags of its own have a significant coefficient. For stock market returns, only the first lag of returns and the dummy variable for September 2001 are statistically significant at conventional levels. Finally, lagged stock returns, the dummy variable and the exogenous variable that measures the expectations about the policy rate have significant parameters in the equation for the federal funds funds rate. Nevertheless, even if individual parameters are insignificant in an equation, they may be jointly statistically significant. For this reason, F-statistic and its p-value are reported for each equation in the system. The null hypothesis that all coefficients on the regressors are equal to zero can be rejected in each case. The adjusted R-squared value varies from 0.092 for the stock market returns to 0.998 for the federal funds rate. Thus, the model overall appears to be specified well.

With respect to the identified structural model, the results are summarised in Figures 3.7 and 3.8. Generally, the impulse response functions are overall similar to those reported in Section 3.8.3. As Figure 3.7 indicates, an unexpected 1-percentage-point increase in the funds rate has a negative effect on inflation and output gap that becomes statistically significant after several periods. There is no evidence of the price puzzle since inflation have fallen by 0.6 percentage points within ten months following the shock and remained there until the end of horizon. Thus, monetary policy impact on inflation seems stronger as compared to the full sample (1989:2 – 2007:7) analysis using the augmented SVAR model. The decline in output gap after five months is around 0.6 percentage points but it is less persistent than in the case of inflation and is reversed in approximately one year following the shock. Overall, the model produces the dynamics of macroeconomic variables in response to a positive interest rate shock that are in line with predictions by standard economic theories and the transmission mechanism of monetary policy.

Figure 3.7: Impulse responses to FFR shock - augmented model (since February 1994)

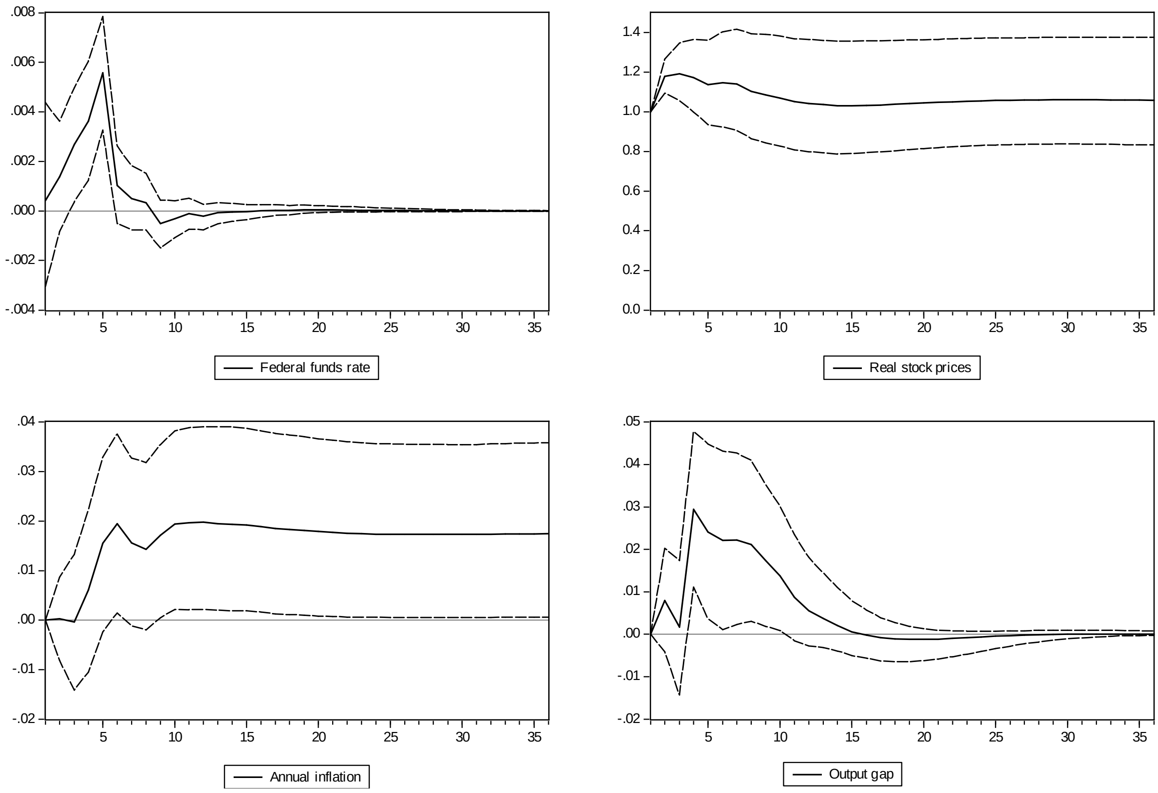


Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

As expected, unexpected monetary policy tightening also has a negative and statistically significant effect on real stock prices that works through the corresponding policy impact on either future expected cash flows (dividends) or the discount rate, i.e. a risk-free rate and risk premium, or the combination of the two. The immediate decline of 7.61% is much more pronounced than 3.53% drop reported for the sample period starting in 1989:2. A typical change in the funds rate target of 0.25% leads to a 1.90% decrease in real aggregate stock prices. This also implies that monetary policy effects on the stock market grew stronger since 1994. The stock market response is statistically significant for two months and is almost entirely reversed thereafter, indicating a quick adjustment of financial markets to new information.

Figure 3.8 demonstrates the dynamic effects of a 1% positive stock price shock. Both inflation and output respond positively and significantly, albeit the response becomes significant with a delay and the magnitude of the impact is moderate. With respect to the monetary policy reaction to the stock market, the federal funds rate increases only a little initially and the response is insignificant. The policy reaction becomes statistically significant after several months with the funds rate having increased by approximately 0.6 of a basis point in the fifth month (at the peak) that is an even small increase than is found in the full sample analysis. Thereafter, it quickly falls back to the initial level. Following an unexpected increase in aggregate real stock prices by 10%, the funds rate would rise by 6 basis points over the five-month period. As compared to the baseline results where the funds rate would rise by 90 basis points over the same time period, the reaction is almost five times smaller. Hence, the Fed's response to stock market developments does not appear to be quantitatively and economically important once two forward-looking variables are included in the VAR system. It is likely, that the additional variables, especially the expected level of the policy rate, incorporate similar information relevant for policymakers as stock prices do.

Figure 3.8: Impulse responses to SP shock - augmented model (since February 1994)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure 3.7 notes.

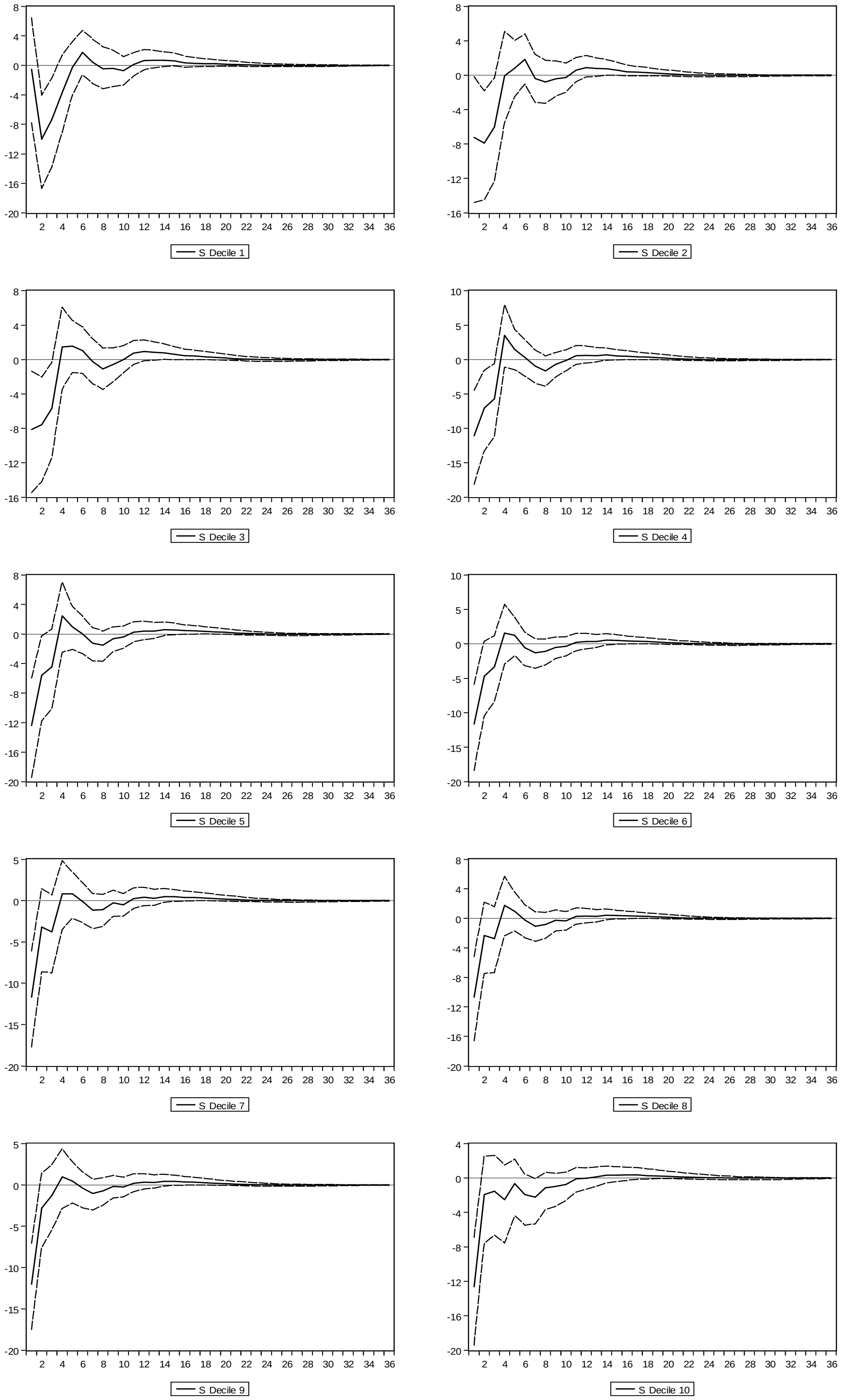
In sum, the results for the period 1994:2 – 2007:7 indicate that contractionary monetary policy depresses economic activity and leads to a significant decline in the stock market. On the other hand, there is little evidence of a strong simultaneous response by the Fed to developments in the stock market. The funds rate response is somewhat delayed and rather small in magnitude.

3.9.2 Size-sorted portfolios

With respect to inflation and output gap, the impulse response functions obtained using stock portfolio returns are very similar to the IRFs reported using the stock market returns. Consequently, in this and the following sections the focus is on the responses of real stock returns across various stock portfolios to a monetary policy shock and the federal funds rate response to a stock price shock.

According to the credit channel of monetary policy transmission, tighter policy stance increases the external finance premium leading to lower stock prices. There are two mechanisms that explain how monetary policy can have the impact on this premium. The bank-lending channel works via changes in the overall supply of intermediate credit. In response to contractionary monetary policy, banks reduce their lending and charge higher rates of interest, i.e. funds raised externally become more expensive. Consequently, more bank-dependent firms reduce their investment spending and their stock prices fall due to lower expected future net cash flows (Kashyap, Stein and Wilcox, 1993). Alternatively, the balance-sheet channel works via changes in firm's creditworthiness. Following monetary policy tightening, the balance sheet position of a firm deteriorates and the external finance premium rises. Firstly, monetary policy directly increases debt-servicing expenses and reduces the value of assets that serve as a collateral to obtain external financing (Bernanke and Gertler, 1995). This affects investment and spending decisions of a firm. Secondly, it may also affect firm's net worth indirectly through its negative impact on overall economic activity and sales revenue, i.e. expected future net cash flows. Overall, the negative effects of monetary policy tightening on stock prices are expected to be stronger for relatively financially constrained firms.

Figure 3.9: Impulse responses of real stock prices to FFR shock - augmented model (since February 1994; size decile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across ten size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

Since small firms are typically viewed as more financially contrained and more heavily reliant on bank lending than large firms, smaller stocks should decline more than larger stocks in response to a positive interest rate shock. As shown in Figure 3.9, an unexpected 1-percentage-point hike in the funds rate leads to a sharp decline in real stock prices across ten size-sorted portfolios, where the tenth decile denotes the largest stocks. With an exception of the smallest stocks, the immediate monetary policy impact is statistically significant and economically meaningful across individual portfolios. Panel A of Table 3.3 (first column) reports the initial-period impulse responses of stock prices for ten size portfolios.

Table 3.3: Impulse responses of real stock prices to FFR shock - augmented model (since February 1994; size and value decile portfolios)

	Panel A: Size portfolios		**Panel B: Value portfolios**	
Decile	**Initial period**	**Second period**	**Initial period**	**Second period**
1st	-0.514 (-7.800; 6.470)	-9.973* (-16.712; -4.034)	-11.843* (-19.644; -4.838)	-1.953 (-8.409; 3.306)
2nd	-7.250* (-14.800; -0.182)	-7.907* (-14.489; -1.817)	-10.981* (-16.743; -5.859)	-3.220 (-7.983; 0.908)
3rd	-8.156* (-15.448; -1.350)	-7.604* (-14.223; -2.028)	-13.288* (-19.720; -7.861)	-3.677 (-9.420; 0.879)
4th	-11.068* (-18.149; -4.471)	-7.073* (-13.325; -1.581)	-16.442* (-21.681; -11.649)	-2.187 (-6.946; 2.067)
5th	-12.392* (-19.449; -5.992)	-5.605* (-11.783; -0.238)	-13.671* (-18.786; -9.090)	-1.183 (-5.803; 2.867)
6th	-11.652* (-18.369; -5.881)	-4.672 (-10.482; 0.389)	-14.364* (-20.268; -9.269)	-3.204 (-8.855; 1.134)
7th	-11.694* (-17.709; -6.097)	-3.189 (-8.622; 1.440)	-14.797* (-19.892; -10.376)	-2.266 (-6.843; 1.658)
8th	-10.670* (-16.586; -5.198)	-2.335 (-7.484; 2.193)	-13.782* (-18.269; -9.573)	-2.970 (-7.158; 0.810)
9th	-12.006* (-17.473; -7.067)	-2.798 (-7.575; 1.397)	-14.927* (-20.070; -10.340)	-4.492* (-9.360; -0.181)
10th	-12.601* (-19.419; -6.879)	-1.937 (-7.632; 2.512)	-9.596* (-14.797; -4.621)	-4.399 (-9.472; 0.225)
Spread	7.875* (1.719; 14.574)	-11.143* (-16.470; -5.717)	1.796 (-4.076; 8.528)	-1.636 (-6.382; 3.971)

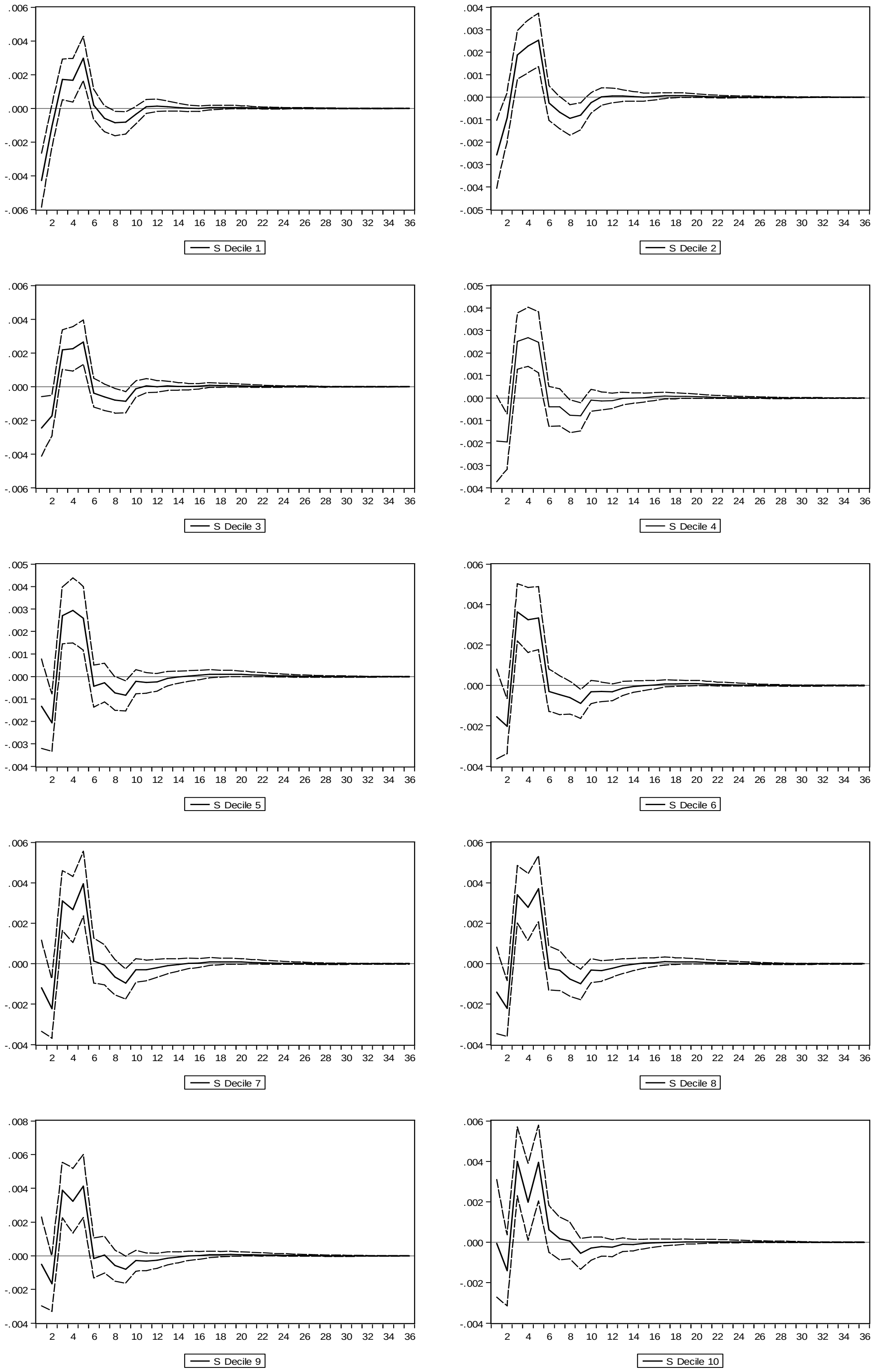
Notes: This table reports the central measures (median) of the initial and second-period impulse responses of real stock prices following a 1-percentage-point contractionary monetary policy shock. Panel A reports the initial- (first column) and second-period (second column) impulse responses of stocks across ten size-sorted portfolios. Panel B reports the initial- (first column) and second-period (second column) impulse responses of stocks across ten value-sorted portfolios. The augmented SVAR model is estimated for each decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^a_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the monthly real stock returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015) are presented in parentheses. * denote statistically significant impulse responses, i.e. the probability bands do not include zero.

Typically, the magnitude of the decline is even greater than that of the aggregate stock market reported in the previous section. As a result of monetary policy tightening, the first decile stocks fall by 0.51% on the impact, while stocks in the tenth decile portfolio are depressed 12.60% below the pre-shock level. Furthermore, the spread return, i.e. the return on the smallest portfolio minus the return on the largest portfolio, responds positively and significantly to an unexpected increase in the funds rate. This implies that monetary policy has a significantly stronger initial effect on the largest stocks as compared to the smallest stocks. Across the remaining size portfolios, the negative initial impact of monetary policy tends to increase in magnitude as the firm size increases; however, the change is not monotonic.

The intial-period responses do not provide support for the credit channel theory. On the other hand, the dynamics of monetary policy effects reveal several interesting findings. First of all, Figure 3.9 implies that the response of stocks in lower deciles (first to fifth) is more persistent as compared to larger stocks in the sixth to tenth deciles. The second-period responses to the monetary policy shock are summarised in the second column of Panel A (Table 3.3). The smallest stocks decrease further, i.e. to 9.97% below the pre-shock level. In contrast, the initial decline in the tenth decile portfolio is mostly reversed as they are only 1.94% lower in the second month after the shock. Furthermore, the magnitude of the response decreases almost monotonically as the firm size increases. Secondly, the monetary policy effect on smaller firms in the bottom five deciles remains statistically significant for several periods following the shock. Conversely, it is not statistically significant beyond the initial period for larger firms in the top five deciles. Moreover, the response of the spread return is now negative and statistically significant indicating that the smallest stocks decline significantly more as compared to the largest stocks. Hence, the second-period responses imply that smaller firms have greater exposure to the monetary policy risk and their stocks are more sensitive to changes in interest rates as compared to larger firms. This is consistent with the credit channel of monetary policy transmission. Overall, the evidence indicates the “delayed size effect” of monetary policy.

The impulse responses of the federal funds rate to a 1% stock price shock for each size portfolio are presented in Figure 3.10. There is some tendency for the funds rate to decline initially and for some portfolios this counter-intuitive negative response is also statistically significant. This may imply that the Fed accommodates stock price increases by cutting its policy rate. As it has been found previously for the stock market, there is the evidence of a delayed positive reaction of monetary policy to stock price developments that eventually also becomes statistically significant. Nevertheless, this reaction again appears to be very small and does not appear to be quantitatively important.

Figure 3.10: Impulse responses of the federal funds rate to SP shock - augmented model (since February 1994; size decile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across ten size-sorted portfolios. The augmented SVAR model is estimated for each size decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

To sum up, contractionary monetary policy has a strong, negative and economically significant effect on the valuation of stock portfolios sorted by the firm's size. On the other hand, stock price developments do not appear to be of much importance to monetary policymakers. The initial-period impulse responses indicate that larger firms are more sensitive to monetary policy shocks than smaller firms. However, the opposite is true in the second period. Overall, the results provide some evidence of the credit channel of monetary policy transmission. As small firms tend to be more bank-dependent and typically have weaker relationships with financial intermediaries, the finding that monetary policy tightening depresses smaller stocks more than larger stocks (even though with a delay of one month) is indicative of the bank lending channel (Kontonikas and Kostakis, 2013).

To some extent, the results in this section are in line with similar studies, albeit those studies report the evidence in favour of the credit channel with respect to an immediate monetary policy effect. For instance, Tsai (2011) provides the empirical evidence that small firms are more exposed to the monetary policy-related risk than larger firms over the period 1984 – 2008. Similarly, Maio (2014) demonstrates that small stocks decline by more in response to a contractionary monetary policy shock than large stocks in the post-1983 and post-1993 periods. Nevertheless, the monetary policy shock does not have a statistically significant impact on returns across ten size portfolios. Also, the results are mixed with respect to whether the differential responses across portfolios are statistically significant (Maio, 2014). Kontonikas and Kostakis (2013) do not find any significant differential effect of monetary policy across size-sorted stock portfolios in the period 1983 – 2007.[166] Similarly to Maio (2014), they also show that there is no statistically significant monetary policy impact on individual stock portfolio returns. On the contrary, the analysis here shows that there is an immediate, strong and statistically significant monetary policy impact on real stock prices across ten size portfolios in the post-1993 period. Also, while larger firms are more sensitive to monetary policy shocks than smaller firms initially, the policy effect on smaller stocks becomes significantly stronger as compared to larger stocks in the second period.[167]

The significant delayed effect of monetary policy on smaller stocks could be explained in several ways. Firstly, stocks of smaller firms are typically much less liquid relative to large stocks (Amihud, 2002; Jensen and Moorman, 2010). Thus, investors may

[166] They do not consider the dynamic responses of ten size portfolios for this sample period.

[167] The statistical significance may be due to the implementation of a more appropriate identification of the monetary policy shocks in this chapter. Kontonikas and Kostakis (2013) apply the generalised impulse response approach, while the results reported in Maio (2014) are based on the simple monthly regressions of portfolio returns.

be able to respond to monetary policy news instantly by trading the largest, i.e. the most liquid, stocks. With respect to smaller and relatively illiquid stocks, it may take more time for their prices to fully reflect the impact of monetary policy shocks.[168] This argument is also consistent with the findings that larger stocks are associated with greater trading activity, whilst smaller stocks are traded much less frequently (James and Edmister, 1983; Chordia, Huh and Subrahmanyam, 2007). The study by Longstaff (2009) may also provide some explanation. The possibility of non-marketability of some assets, i.e. they cannot be readily sold and/or bought, has implications for the portfolio choice of investors and trading volumes. The model shows that a less patient investor ends up having a much higher share of the liquid asset in the investment portfolio, while a more patient investor holds more of the illiquid asset. Also, the existence of illiquid assets is associated with an increase in trading activity of the liquid asset and a decrease in trading of the illiquid one.

Alternatively, the finding of the stronger and more instantaneous monetary policy impact on larger stocks could possibly be explained by either the liquidity pull-back or portfolio rebalancing hypotheses (Nyborg and Ostberg, 2014). Provided that monetary policy affects the interbank market for liquidity, tighter liquidity conditions following a contractionary monetary policy shock may trigger asset sell-off by banks and levered investors with the strongest effect felt on the most liquid stocks. With respect to the portfolio rebalancing, investors may also be reducing their exposure to stocks due to unexpectedly higher market-wide uncertainty. As the most liquid stocks are easier and less costly to sell quickly, trading would potentially be concentrated in this market segment. Finally, the learning process of an investor with information capacity (or attention) constraints could possibly offer another explanation. For instance, Peng (2005) argues that larger stocks are likely to receive more attention from investors, thus, greater capacity allocation and larger supply of information lead to quicker stock price adjustments to fundamental shocks.

3.9.3 Value-sorted portfolios

As it have been explained earlier, value firms are likely to be relatively financially constrained and more heavily dependent on cash flows as they are characterised by high

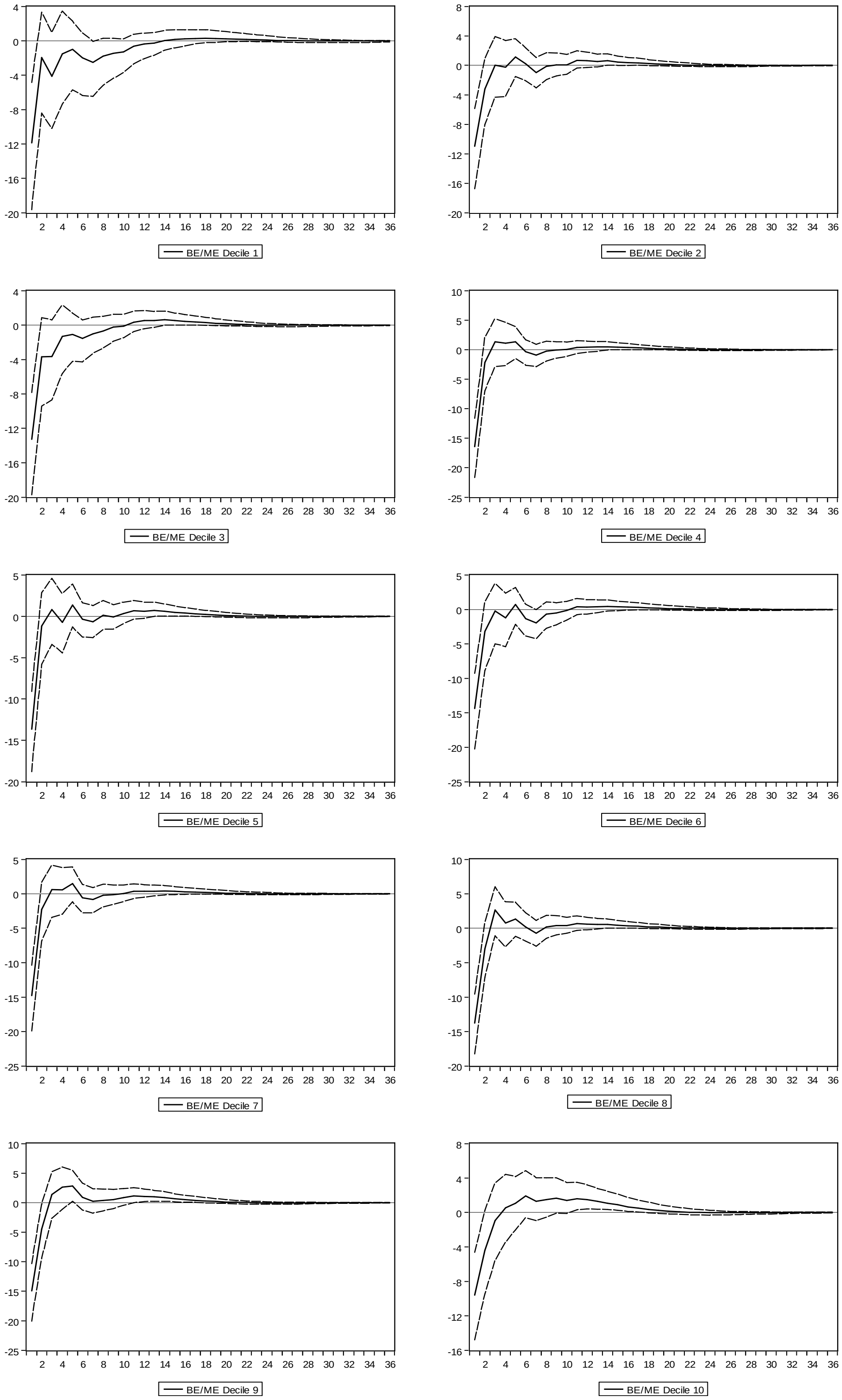
[168] With respect to daily event studies, Cenesizoglu (2011) reports that larger stocks decline by more than smaller stocks on the days of the federal funds target rate announcements. For the UK, Florackis, Kontonikas and Kostakis (2014) demonstrate that monetary-policy-induced macro-liquidity shocks have stronger effects on portfolio returns of more liquid stocks, possibly due to changes in aggregate stock market liquidity conditions. To this respect, Goyenko and Ukhov (2009), provide the evidence that US monetary policy shocks have a significant impact on stock market illiquidity with tighter policy increasing stock illiquidity.

current and near-future earnings and dividends, while they have poor growth opportunities. Accordingly, value stocks tend to be undervalued in the market and also are likely to be more sensitive to the interest rate risk as compared to growth stocks (Fama and French, 1995; Kontonikas and Kostakis, 2013, Lioui and Maio, 2014). Thus, stock portfolios sorted by book-to-market value ratio may shed some light into the balance sheet mechanism of the credit channel.

Figure 3.11 shows the dynamic effects of a contractionary monetary policy shock on stocks across ten value portfolios. The immediate response of real stock prices is always negative and statistically significant but it is almost never significant beyond the initial period. The magnitude of the responses is also very large indicating that monetary policy have economically important effects on real stock prices. Comparing the reaction of the extreme deciles, the most value stocks (the tenth decile) and the most growth stocks (the first decile) fall by 9.60% and 11.84%, respectively. This is not in line with the predictions of the credit channel theory. Nevertheless, the response of the spread return between these deciles is statistically insignificant. Thus, the monetary policy impact differential on value versus growth stocks is not statistically significant. As reported in Table 3.3 Panel B, there is no clear trend in the relationship between the stock price response to the policy shock and book-to-market value ratio. For example, while stocks in the ninth decile are more responsive than stocks in the second decile, seemingly providing some support for the credit channel, the greatest decline is that for the stocks in the fourth decile portfolio. Meanwhile, the responses of the third, fifth and eight deciles are very similar.

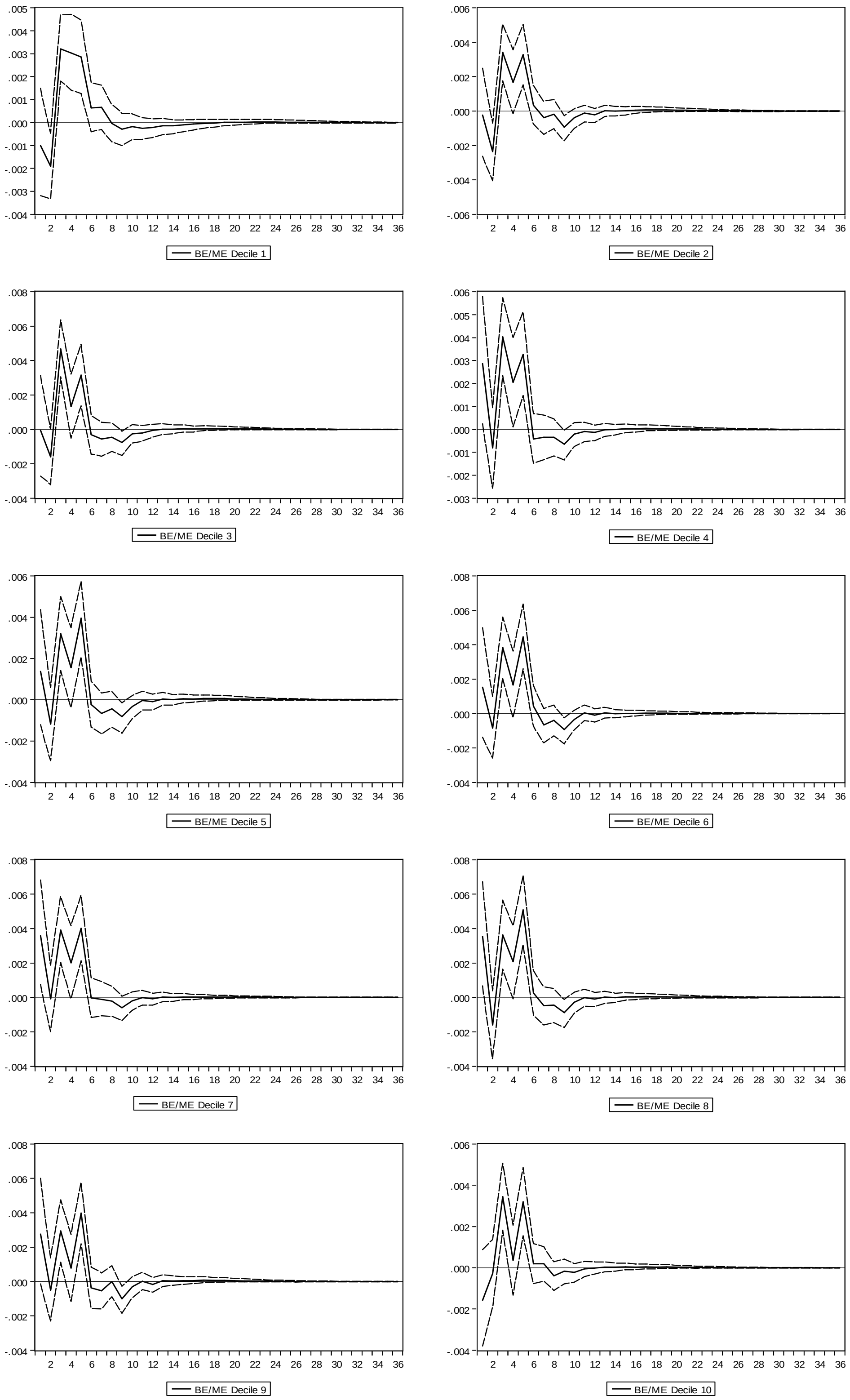
Unlike with some size portfolios, there is little evidence of any persistence in the effect of monetary policy on stock prices. Across all ten portfolios, the initial decline is largely reversed in the next month. As can be noted in the second column in Panel B of Table 3.3, the spread return declines in the second period indicating that the most value stocks are more negatively affected by a monetary policy shock than the most growth stocks. Nevertheless, this differential impact is again statistically insignificant. Across the remaining portfolios, the second-period response does not change monotonically as one moves from lower to higher deciles. The smallest decline can be noted in the fifth decile, alhough it is insignificant. On the other hand, stocks in the ninth decile decline the most and it is the only case where the response is also statistically significant.

Figure 3.11: Impulse responses of real stock prices to FFR shock - augmented model (since February 1994; value decile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across ten value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each value decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

Figure 3.12: Impulse responses of the federal funds rate to SP shock - augmented model (since February 1994; value decile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across ten value-sorted portfolios. The augmented SVAR model is estimated for each value decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

Figure 3.12 shows the impulse responses for the federal funds rate following a positive stock price shock. There is a delayed positive and significant reaction; however, the absolute size of the response is very small as in all previous cases. Thus, these results are very similar to those reported for the size portfolios and for the stock market indicating the absence of the strong policy reaction to stock price developments.

Overall, the findings do not provide the support for the differential impact of monetary policy shocks on value versus growth stocks and, thus, are inconsistent with the credit channel of monetary policy transmission. Other similar studies also fail to show that value stocks are affected significantly more by monetary policy shocks as compared to growth stocks in the more recent period (Kontonikas and Kostakis, 2013; Maio, 2014). Also, they do not find the evidence that monetary policy effect on individual portfolios is statistically significant. On the other hand, the analysis here indicates that monetary policy has a strong, negative and statistically significant impact on stock prices across ten value-sorted portfolios.

3.9.4 Double-sorted size-value portfolios

The results obtained using either size-sorted or value-sorted stock portfolios provide some insight into the size and value effects of monetary policy, respectively. Nevertheless, such univariate sorts only control for one characteristic at a time. For instance, size portfolios do not account for the BE/ME ratio of firms and, similarly, value portfolios do not take into account the size of firms. Consequently, this section investigates the monetary policy interaction with stock returns on double-sorted portfolios that are formed on the basis of both the size and the BE/ME ratio of firms.

Table 3.4 Panel A reports the initial-period impulse responses across the twenty five double-sorted size-value portfolios following a contractionary monetary policy shock. The immediate stock price response is always negative and quantitatively important. It is also mostly statistically significant with an exception of several smaller stock portfolios. In the case of the smallest quintile, it is only the most value stocks that are significantly affected, for instance. With respect to the size effect of monetary policy, for all value quintiles, larger stocks are generally more responsive than smaller stocks with the magnitude of the decline increasing when moving from the first to the fifth size quintile. Nevertheless, the relationship between the degree of responsiveness and the size is not strictly monotonic. This counter-intuitive differential impact of monetary policy on the smallest versus largest stocks is least pronounced for the most value stocks (fifth quintile).

With respect to the bottom four value quintiles, the difference is typically above 10% (percentage points) as compared to only around 4% difference for the most value portfolios. Within the first value quintile (most growth stocks), following an unexpected 1-percentage-point increase in the federal funds rate, the largest stocks decline by 11.36% as compared to a 3.06% drop in the smallest stocks. For the stocks in the fifth value quintile, the largest and smallest stocks are 13.78% and 9.70% lower, respectively. Overall, the smallest stocks become more sensitive to monetary policy shocks when they are also the most value stocks.

Table 3.4: Impulse responses of real stock prices to FFR shock - augmented model (since February 1994; double-sorted size-value quintile portfolios)

Panel A: Initial period					
	1st Value quintile	*2nd Value quintile*	*3rd Value quintile*	*4th Value quintile*	*5th Value quintile*
1st Size quintile	-3.061 (-14.114; 7.613)	-1.489 (-9.729; 6.705)	-3.819 (-10.467; 2.407)	-4.894 (-10.572; 0.586)	-9.704* (-15.560; -4.116)
2nd Size quintile	-7.145 (-16.801; 2.244)	-11.376* (-18.033; -5.123)	-11.440* (-17.128; -6.396)	-12.631* (-18.779; -7.133)	-12.171* (-18.556; -6.199)
3rd Size quintile	-10.176* (-20.031; -1.343)	-14.271* (-20.527; -8.566)	-14.726* (-20.245; -9.842)	-15.689* (-21.208; -10.769)	-14.252* (-19.748; -9.421)
4th Size quintile	-8.436* (-16.663; -0.627)	-15.275* (-20.896; -10.196)	-16.295* (-21.775; -11.607)	-14.394* (-19.144; -9.832)	-15.869* (-21.211; -10.961)
5th Size quintile	-11.356* (-18.396; -5.241)	-14.476* (-20.369; -9.521)	-14.417* (-20.251; -9.471)	-16.120* (-21.412; -11.598)	-13.785* (-19.606; -8.338)
Panel B: Second period					
	1st Value quintile	*2nd Value quintile*	*3rd Value quintile*	*4th Value quintile*	*5th Value quintile*
1st Size quintile	-11.663* (-21.666; -2.980)	-8.581* (-15.733; -2.148)	-7.296* (-12.850; -2.295)	-7.577* (-12.533; -3.110)	-9.433* (-15.693; -4.327)
2nd Size quintile	-9.266* (-17.690; -1.815)	-6.253* (-12.248; -1.101)	-4.816* (-9.943; -0.312)	-7.871* (-13.745; -2.945)	-8.727* (-14.881; -3.490)
3rd Size quintile	-8.559* (-16.780; -1.370)	-3.581 (-9.318; 1.406)	-2.918 (-8.043; 1.379)	-4.802* (-9.860; -0.342)	-3.915 (-8.964; 0.373)
4th Size quintile	-3.929 (-10.600; 2.157)	-2.256 (-7.437; 2.134)	-1.847 (-7.184; 2.601)	-1.024 (-5.514; 3.043)	-4.875* (-9.709; -0.644)
5th Size quintile	-2.012 (-7.753; 2.668)	-2.252 (-7.174; 1.988)	-1.978 (-7.093; 2.263)	-0.541 (-5.013; 3.425)	-3.136 (-8.613; 1.907)

Notes: This table reports the central measures (median) of the initial- and second-period impulse responses of real stock prices following a 1-percentage-point contractionary monetary policy shock. Panel A reports the initial-period and Panel B reports the second-period impulse responses of stocks across twenty five double-sorted size-value stock portfolios. The augmented SVAR model is estimated for each decile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Table 3.3 notes.

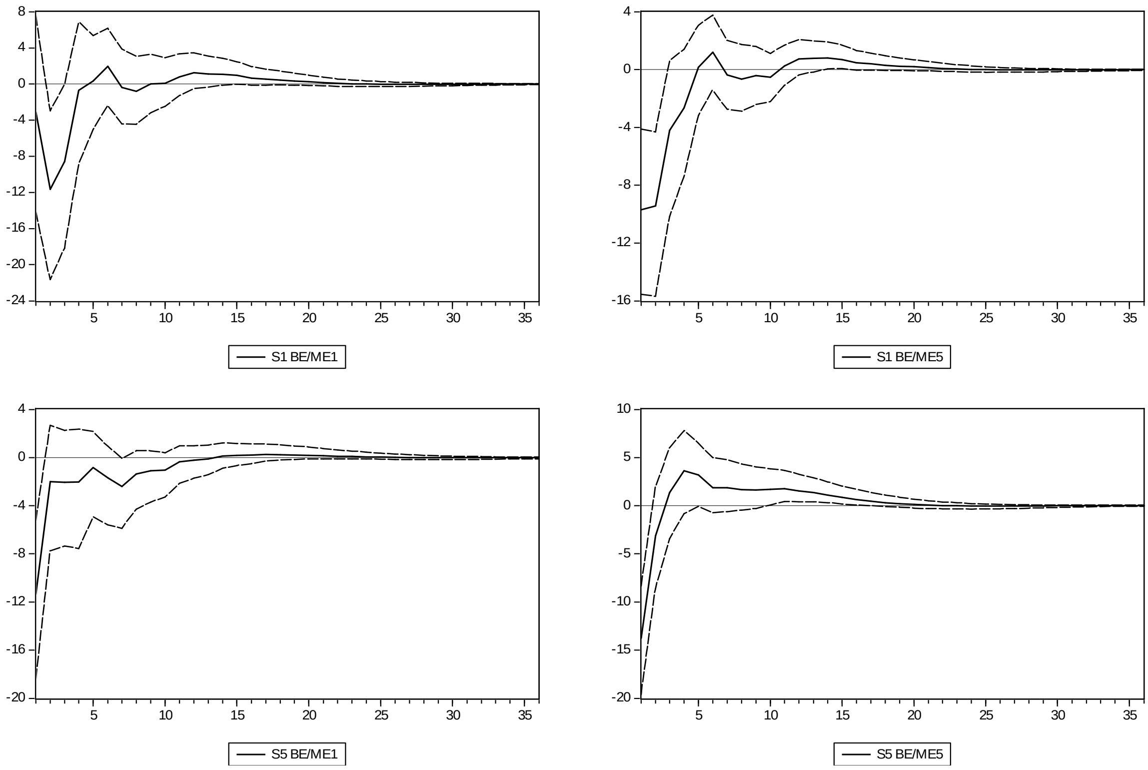
Second, monetary policy always has a stronger immediate effect on the most value stocks than the most growth stocks. Within each size quintile, stocks with the highest BE/ME ratio always decline more than stocks with the lowest BE/ME.[169] In addition, the monetary policy impact differential on the most value versus most growth stocks is larger for the portfolios of the smallest stocks (6.64%) as compared to the portfolios of the largest stocks (2.42%). In the case of the largest size quintile, this differential is the smallest of all. On the other hand, there is no monotonic increase in the magnitude of the effect across the value quintiles with an expection of the smallest stock portfolios. Furthermore, the magnitudes of the responses are quite similar if the second, third and fourth value portfolios are compared for each size quintile. Overall, the most value firms appear to be more exposed to monetary policy risk than the most growth firms, in line with the credit channel. This is especially true when these stocks are also smaller stocks. Meanwhile, the evidence for the differential monetary policy impact on value versus growth firms is far weaker with respect to the middle value quintiles for each size quintile.

Panel B of Table 3.4 provides the results with respect to the second-period monetary policy effect. The reaction of real stock prices is always negative; however, it is mostly smaller stocks in the first two size quintiles that experience a statistically significant decline. Meanwhile, the impact on larger stocks in the top two size quintiles is typically insignificant. In contrast to Panel A, there is strong evidence of the size effect within each value quintile.[170] For instance, the smallest most value stocks decline by 9.43%, whilst the largest most value stocks only decrease by 3.14% in the next period after unexpected monetary policy tightening. This confirms the delayed size effect with respect to monetary policy shock as previously reported for ten size-sorted portfolios. Furthermore, the absolute magnitude of the response tends to gradually decline as the firm size increases, albeit it is not strictly monotonous for some value quintiles. On the other hand, the value effect seems to diminish or even reverse in the second period following the shock. For each size quintile, there is some decline in the response when moving away from the most growth portfolio to the third quintile and subsequently tends to increase slightly for the remaining value portoflios. Nevertheless, in the case of smaller stocks, the most value stocks are less sensitive to an interest rate shock than the most growth stocks.

[169] Nevertheless, the response of the spread return between the most value and the most growth stocks for each size quintile is not statistically significant. Thus, the differential impact on value versus growth stocks does not appear to be significant. On the other hand, the difference between the response coefficients of the extreme quintiles is much greater than in the case of decile value portoflios.

[170] Based on the response of the spread return between the smallest and the largest stocks, the differential impact seems to be statistically significant.

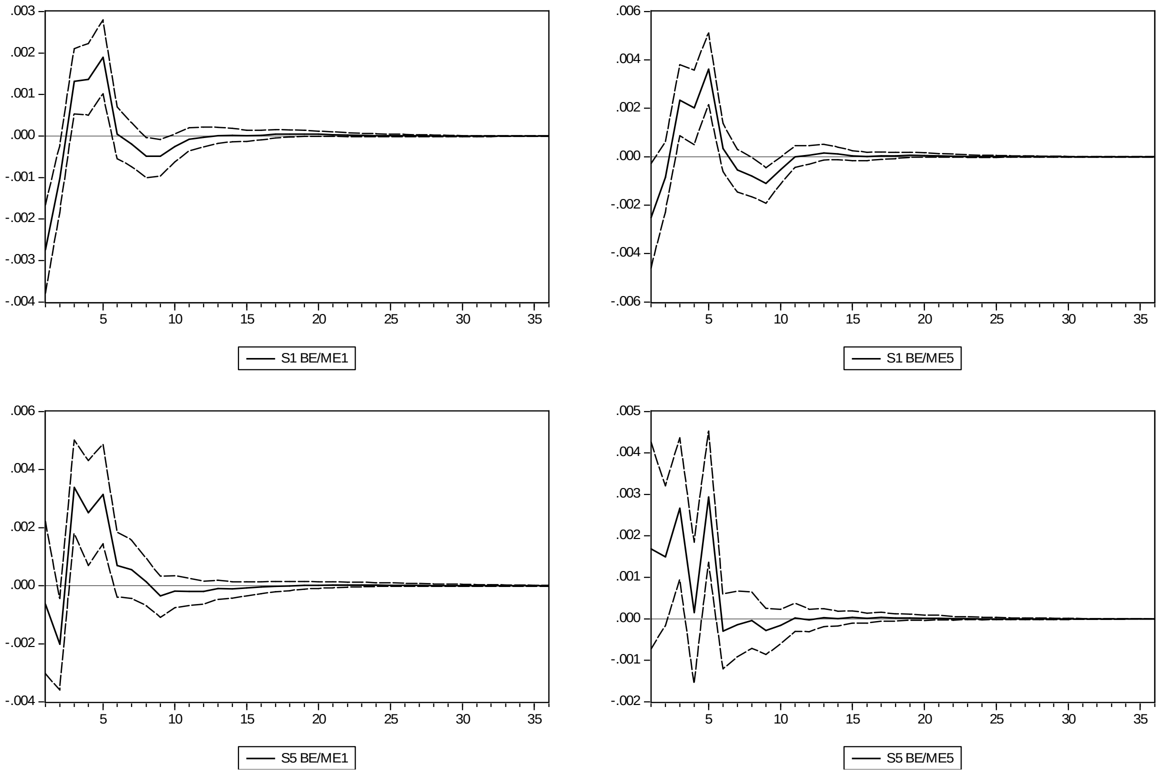
Figure 3.13: Impulse responses of real stock prices to FFR shock - augmented model (since February 1994; size-value quintile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

Figure 3.13 plots the responses of real stock prices for the four extreme size-value portfolios. The selected portfolios represent the intersection of the first and the fifth size quintiles with the first and the fifth value quintiles. The top row in Figure 3.13 compares monetary policy effects on stock prices for the most growth (left) and the most value (right) portfolios within the smallest size quintile. The bottom row shows the stock price response of the most growth and the most value portfolios for the largest stocks. The first and second columns clearly indicate the existence of the delayed size effect of monetary policy. The decline in the stock prices of the smallest firms is also more persistent as compared to a decline in the largest stocks. Also, the response of the smallest stocks is statistically significant for several periods after the shock, while the largest stocks affected significantly only on the impact. The value effect is also evident, but it is only present in the initial period, i.e. monetary policy has an immediate differential effect on value versus growth firms, with value stocks being affected more adversely.

Figure 3.14: Impulse responses of the federal funds rate to SP shock - augmented model (since February 1994; size-value quintile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate (i_t) following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing the real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure 3.7 notes.

Finally, Figure 3.14 depicts the dynamic response of the federal funds rate following a positive stock price shock with respect to the four selected portfolios. In all cases, the response eventually becomes positive and significant. The initial negative and significant reaction of the policy rate in the case of the smallest stocks is a bit puzzling. On the other hand, the absolute size of the impulse responses are again very small. This offers only weak evidence for the strong and quantitatively important policy reaction in response to stock price developments.

Overall, the empirical evidence implies that monetary policy shocks have a stronger impact on small stocks that are likely to be more financial constrained and more bank-dependent as compared to large firms. However, this size effect only becomes evident with the delay of one month. After controlling for the firm's size, the results also indicate the immediate value effect of a monetary policy shock, suggesting that the balance sheet mechanism of the credit channel may be active.[171] Thus, the results obtained using double-

[171] This interpretation of the differential monetary policy impact on the value versus growth firms is also considered by Kontonikas and Kostakis (2013).

sorted portfolios provide some support for the credit channel of monetary policy transmission. As previously, there is no evidence in favour of the strong monetary policy response to stock price developments.

3.10 Robustness analysis

This section examines the robustness of the main empirical findings from the estimation of the augmented SVAR model in Section 3.9. The analysis framework is altered in a number of ways. First, alternative stock portfolios are used in the SVAR model. Second, an alternative lag length is chosen. Third, an alternative measure of output gap is considered. Fourth, the exogenous dummy variable for September 2001 is excluded from the model. Fifth, the model is estimated over the sample period that extends until December 2008. In addition, an alternative data transformation is used and inflation variable is included in the state vector without differencing. Finally, the dummy variable that takes value 1 during the US recession periods is included as an exogenous variable. In order to preserve space, the results are reported for the stock market returns, the extreme deciles of the size and value portfolios, and four double-sorted portfolios representing the intersections between the extreme size and value quintiles. All tables and figures are provided in the Appendix. Overall, the main empirical findings are reasonably robust to all alternations to the augmented model.

3.10.1 Alternative stock portfolios (quintile portfolios)

Instead of decile stock portfolios, the univariate sorts of quintile portfolios formed on the market value (size) and BE/ME ratio are used. The results are reported in Figures A3.3 – A3.6 and Table A3.2. With respect to the firm’s size, an unexpected monetary policy tightening depresses real stock prices across the board with the initial impact being stronger for larger firms. Generally, the immediate stock price response is always statistically significant. Nevertheless, the response of the spread return is statistically insignificant indicating that the differential impact is not significant. In line with the main results, there is strong evidence of the size effect in the second period following the monetary policy shock. The smallest firms are significantly more sensitive to the shock as compared to the largest stocks. With respect to value portfolios, real stock prices always decline immediately in response to an unexpected increase in the federal funds rate. The initial response is also statistically significant for all five value quintiles. Initially, the most value stocks are slightly more negatively affected than the most growth stocks. However, the response of the spread return is insignificant. The same is true for the second-period impulse responses, although the effect of monetary policy is now insignificant for all portfolios. In terms of a stock price shock, there is no convincing evidence for an economically significant monetary policy reaction to stock developments. Overall, the results are qualitatively similar to those obtained in the main analysis.

3.10.2 Alternative stock portfolios (industry portfolios)

The main model is now estimated for ten industries. Monthly stock returns (excluding dividends) on value-weighted portfolios are provided by Kenneth R. French. The ten industries include consumer durable goods, consumer non-durable goods, manufacturing, energy, high technology, telecommunications, wholesale/retail, healthcare, utilities and "other" industries. The results are summarised in Figures A3.7 – A3.8. Across the industries, real stock returns decline initially and bounce back over the next several periods in response to monetary tightening. Typically, the response is statistically significant in the first month only. The most affected are durable consumer goods and manufacturing industries with stock prices declining by 17.71% and 16.98% on the impact, respectively. In addition, retail/wholesale stocks are also very sensitive to changes in monetary conditions, while energy and utilities sectors are less negatively affected. The most resilient are healthcare and high-tech stocks since the monetary policy impact on these stocks is statistically insignificant, albeit still negative.

The results are broadly in line with the existing empirical work.[172] For instance, strong and significant monetary policy effects on durables and retail industry are reported in several other studies (Erhman and Fratzscher, 2004; Bernanke and Kuttner, 2005; Kurov, 2010). Jansen and Tsai (2010) find a strong and negative response of manufacturing sector to monetary policy shocks. Small or moderate monetary policy effects on energy, utilities and healthcare sectors are reported by Thorbecke (1997), Bernanke and Kuttner (2005), Kurov (2010). Also, Kontonikas, MacDonald and Saggu (2013) find an insignificant and small monetary policy impact on the high-tech industry. Generally, the federal funds rate increases in response to a positive stock price shock and the policy reaction to higher stock prices is typically statistically significant. The response

[172] Note that the majority of these studies employ an event study to estimate the monetary policy effect on industry returns.

is broadly similar across the industries. Nevertheless, the magnitude of the peak response of the funds rate is even smaller than in the case of the stock market.

3.10.3 Alternative lag length

The SVAR model in the main analysis is estimated with four lags. In this section, the model is estimated using six and, alternatively, two lags. Figures A3.9 –A3.14 report the results obtained from the SVAR(6) model for the stock market and the selected portfolio returns. Figures A3.15 – A3.20 present the impulses responses obtained by estimating the SVAR(2) model.

In general, monetary policy tightening has a negative and significant impact on real stock prices and economic activity. The size effect is noticeable with the delay of one period, while the value effect is more evident after controlling for the firm's size, i.e. when double-sorted size-value portfolios are used. With respect to a stock price shock, the response of the federal funds rate eventually turns positive and significant; however, it is very small in absolute size. Overall, the main results are robust to the alternative orders of the SVAR model.

3.10.4 Alternative measure of output gap

The main empirical analysis calculates the output gap by deducting the output trend level from the actual series using the Hodrick-Prescott filter. This section substitutes this measure with the output gap obtained using the quadratic trend instead.

The results are summarised in Figures A3.21 – A3.26. A contractionary monetary policy shock has a negative and significant effect on the stock market and economic activity, whilst a positive stock price shock leads to higher economic activity and a slightly higher policy rate after several periods. With regards to portfolio returns, monetary policy tightening has a stronger and more persistent effect on smaller stocks with the delay of one month. In addition, there is also evidence in favour of the value effect of monetary policy following the shock. As previously reported, the monetary policy reaction to stock price shock is very moderate. Overall, the main findings do not change substantially if the alternative output gap measure is used.

3.10.5 Stock market crash in 2001

The augmented SVAR model employed in the main empirical analysis is re-estimated without the dummy variable that takes value of one in September 2001 to account for the stock market crash in the US and is zero otherwise. The impulse response functions are provided in Figures A3.27 – A3.32. Overall, the results are almost identical to those reported in Section 3.9.

3.10.6 Alternative sample period

The sample period is extended to include the global financial crisis period until December 2008 when the short-term interest rate reached the zero lower bound. Two additional dummy variables are included as exogenous variables. The first dummy accounts for the bankruptcy of Lehman Brothers and is set to one for September 2008. The second dummy takes value of one in October 2008 when monthly real stock returns fell sharply and the full effect of Lehman's collapse became apparent. Both dummies take values of zero on all the remaining months.

The results are presented in Figures A3.33 – A3.38. Firstly, the model is estimated using real stock market returns. An unexpected increase in the federal funds rate leads to a decline in real stock prices and a fall in annual inflation and output gap. The response of stock prices is statistically significant; however, there is some statistical uncertainty surrounding the impulse responses of macroeconomic variables. A positive stock price shock gives a boost to economic activity and inflation. Also, it leads to a slightly higher funds rate. In line with the main findings, the monetary policy reaction to the stock market is very small in magnitude. With respect to stock portfolio returns, monetary policy tightening typically has a negative and significant effect on real stock prices. The results imply the size effect of monetary policy that materialises in the second period following the shock. As previously, there is no strong evidence that value stocks are more sensitive to monetary policy shocks than growth stocks using the decile portfolios. However, the most value stocks decline more as compared to the most growth stocks following an unexpected increase in the funds rate when controlling for the firm's size. In response to a positive stock price shock, the policy rate tends to increase significantly as in the case of the stock market. Nevertheless, the reaction is again not economically meaningful. Overall, the findings for the extended sample period 1994:2 - 2008:12 are reasonably in line with the main results reported in Section 3.9, albeit there is somewhat more statistical uncertainty surrounding some impulse responses.

3.10.7 Alternative data transformation

This section employs the alternative data transformation of endogenous variables. Instead of taking annual changes in the price level and leading economic index, the monthly changes in these variables are employed here. Also, the measure of the output gap is replaced with monthly growth in the industrial production index. With respect to inflation variable, this specification is in line with Chapter 2. Hence, the state vector includes the following endogenous variables in this order: a (lagged) monthly change in the log leading economic index (Δlei_{t-1}), a monthly change in the log industrial production index (Δip_t), the monthly consumer inflation rate (π_t^m), the monthly real stock returns (Δsp_t), and the federal funds rate (i_t). As previously, the rate on the one-month federal funds futures contract as of the last day of the previous month is included as an exogenous variable together with the dummy variable that takes the value of one in September 2001 and zero otherwise.

Figures A3.39 – A3.44 report the impulse responses from the estimated model with real returns on the stock market and stock portfolios. With respect to the stock market, an increase in the federal funds rate leads to a fall in real stock prices, a decline in monthly inflation and a contraction in industrial output. The immediate monetary policy impact on stock prices is also statistically significant. Meanwhile, an increase in real stock prices is associated with a positive and significant response of output growth and a positive and significant reaction of the monetary policy rate. However, the effects on monthly inflation are not as clear. Finally, the monetary policy response to stock price shock does not appear to be economically significant.

In the case of stock portfolios, the findings again indicate the delayed size effect with respect to monetary policy shocks. In favour of the credit channel, the most value stocks decline by more and the policy impact is more persistent as compared to the most growth stocks when size-value portfolios are used. Finally, the monetary policy response to a positive stock price shock is positive and significant, albeit with some delay, but is also very small in magnitude. Generally, the results remain largely unchanged as compared to the main results.

3.10.8 Alternative data transformation – undifferenced inflation

In all models considered, the annual inflation variable is differenced to stationarity. This is in line with the original specification of the SVAR model presented in Bjornland and Leitemo (2009). Their study provides the basis for the empirical analysis of Chapter 3. Essentially, the price level is differenced twice, however, the price level is only differenced once for the empirical analysis in Chapter 2. While the previous chapter deals with monthly inflation, the current chapter uses annual inflation and this variable is non-stationary for the sample period under investigation. Nevertheless, this inconsistency across chapters is addressed by testing the robustness of the main findings with respect to annual inflation variable that is not differenced to stationarity. Furthermore, the previous section also partially addresses this point by employing a different data transformation where monthly inflation is used instead.

Figures A3.45 – A3.50 report the impulse responses from the estimated model with real returns on the stock market and stock portfolios with respect to the lowest and highest deciles and quintiles. With respect to macroeconomic variables, a contractionary monetary policy shock has a negative and significant impact on annual inflation that gradually dies out within twenty months. Similarly, output gap declines but the response is not statistically significant and output gap returns to its pre-shock level in about one year. In line with the main findings, an increase in the federal funds rate leads to a significant immediate fall in real stock prices by 6.5%. Meanwhile, unexpectedly higher real stock prices are associated with a positive response of both inflation and output gap; however, the reaction is generally insignificant. As previously, the monetary policy response to a stock price shock does not appear to be economically significant, albeit there is some evidence of the delayed positive response of the funds rate.

Regarding stock portfolio returns, the effect of monetary policy tightening on the smallest stocks turns negative with one period delay but is never statistically significant. On the other hand, the largest stocks decline sharply and significantly on the impact but the effect is reversed in the next month following the shock. In line with the main results, during the second month the first decile stocks show a larger decline as compared to the tenth decile. The most value and the most growth stocks both fall significantly in response to a 1% increase in the federal funds rate, nevertheless, the most growth stocks demonstrate somewhat greater decline. With respect to double-sorted portfolios, Figure A3.49 provides some support for the delayed size effect of monetary policy; however, there is no convincing evidence of the value effect. Finally, the policy response to higher stock prices with respect to portfolios is very small but positive and significant, albeit with some delay.

Overall, the results reported in Section 3.9 are relatively robust to a change in inflation variable, although the evidence in favour of the credit channel of monetary policy transmission mechanism becomes weaker.

3.10.9 Recession dummy variable

As discussed briefly in the literature review, some argue that the impact of monetary policy on stock prices depends on the business cycle and the negative stock price reaction may be driven by the periods of bad economic conditions. Thus, instead of having the dummy variable for the stock market crash, the alternative dummy that takes value of one during recessionary periods, i.e. in 2001:3 – 2001:11, is included as an exogenous variable. The generated impulse responses reflect the dynamic effects of the structural shocks during good economic times, i.e. the dummy variable controls for the recessionary months.

The findings are summarised in Figures A3.51 – A3.56. As compared to the main results, the impulse response functions reported here are very similar. Generally, the decline in stock prices is only slightly smaller in magnitude than its counterparts reported in 3.9. In line with the main findings, the value effect is present in the initial period for size –value portfolios, while the size effect only appears in the next period following the monetary policy shock. With respect to monetary policy reaction to a stock price shock, there is little evidence for economically meaningful response. To sum up, the results reported here are very similar to the results when the business cycle is not controlled for. Thus, the main findings are not driven by a potentially stronger monetary policy interaction with stock prices during recessions.

3.11 Conclusion

This chapter provides the in-depth analysis of conventional US monetary policy interaction with stock prices at both market and portfolio levels focusing on the period 1994 – 2007. The choice of the sample period is motivated by significant changes in the Federal Reserve's communications of its policy decisions to the public at that time. The estimated structural VAR model is specified in the spirit of Bjornland and Leitemo (2009) to allow the simultaneous reaction between monetary policy and real stock returns. The original framework is improved by adding two forward-looking variables following the recommendations by Brissimis and Magginas (2006). This approach helps to eliminate the price puzzle and produces a sharper measure of structural monetary policy shocks.

The main empirical findings show that monetary policy tightening has a strong and negative impact on real stock prices. The negative and significant stock price response to monetary policy shocks is also evident across size, value and double-sorted size-value stock portfolios. The portfolio-level results provide some evidence of the higher sensitivity of small and value stocks to monetary policy risk versus large and growth stocks. Nevertheless, the size effect only becomes evident after one month following the shock, whilst the results are more indicative of the value effect only when controlling for the firm's size. The delayed response of smaller stocks to monetary policy shocks could possibly be linked to their relative illiquidity and less frequent trading or explained through the liquidity pull-back and portfolio rebalancing hypotheses. In addition, the learning process of investors may also play a role. Overall, the findings provide some support, albeit not very strong, for the credit channel of monetary policy transmission. On the other hand, the empirical evidence is not supportive of a strong monetary policy response to stock price developments. The response of the federal funds rate to an unexpected increase in real stock prices is typically positive and becomes statistically significant after several months. Nonetheless, the policy reaction does not appear to be economically significant. This finding is in line with the argument by Bernanke and Gertler (1999) that monetary policy should focus on macro-fundamentals, rather than stock market developments, except insofar as they predict changes in relevant economic variables. Finally, the empirical findings are robust with respect to various alterations to the main model, such as alternative stock portfolios, different lag length, the inclusion of the financial crisis period, alternative data transformation, and etc.

In general, this chapter shows that monetary policy news is an important determinant of financial market developments. Due to the role of stock prices in the transmission mechanism, the understanding of monetary policy impact on stock prices may help policymakers to take appropriate decisions in the future. One potential limitation of this chapter is that unconventional monetary policy tools are not considered and, possibly, the analysis could be extended to this regard.

References for Chapter 3

Abouwafia, H.E., and Chambers, M.J., 2015. Monetary policy, exchange rates and stock prices in the Middle East region. *International Review of Financial Analysis*, 37, p.14-28.

Amihud, Y., 2002. Illiquidity and stock returns: cross-section and time-series effects. *Journal of Financial Markets*, 5(1), p.31-56.

Ammer, J., Vega, C., and Wongswan, J., 2010. International transmission of U.S. monetary policy shocks: Evidence from stock prices. *Journal of Money, Credit and Banking*, 42(6), p.179-198.

Aretz, K., Bartram, S.M., and Pope, P.F., 2010. Macroeconomic risks and characteristic-based factor models. *Journal of Banking and Finance*, 34(6), p.1383-1399.

Arshanapalli, B., Fabozzi, F.J., and Nelson, W., 2006. The value, size, and momentum spread during distressed economic periods. *Finance Research Letters*, 3(4), p.244-252.

Banbura, M., Giannone, D., and Reichlin, L., 2010. Large Bayesian vector auto regressions. *Journal of Applied Econometrics*, 25(1), p.71-92.

Banz, R.W., 1981. The relationship between return and market value of common stocks. *Journal of Financial Economics*, 9(1), p.3-18.

Barakchian, S.M., and Crowe, C., 2013. Monetary policy matters: Evidence from new shocks data. *Journal of Monetary Economics*, 60(8), p.950-966.

Basistha, A., and Kurov, A., 2008. Macroeconomic cycles and the stock market's reaction to monetary policy. *Journal of Banking and Finance*, 32(12), p.2606-2616.

Basu, S., 1983. The relationship between earnings' yield, market value, and return for NYSE common stocks. *Journal of Financial Economics*, 12(1), p.129-156.

Baxa, J., Horvath, R., and Vasicek, B., 2013. Time-varying monetary-policy rules and financial stress: Does financial instability matter for monetary policy?. *Journal of Financial Stability*, 9(1), p. 117-138.

Bernanke, B.S., and Blinder, A.S., 1992. The federal funds rate and the channels of monetary transmission. *The American Economic Review*, 82(4), p.901-921.

Bernanke, B.S., Boivin, J., and Eliasz, P., 2005. Measuring the effects of monetary policy: A factor-augmented vector autoregressive (FAVAR) approach. *The Quarterly Journal of Economics*, 120(1), p.387-422.

Bernanke, B.S., and Gertler, M., 1995. Inside the black box: The credit channel of monetary policy transmission. *Journal of Economic Perspectives*, 9(4), p.27-48.

Bernanke, B.S., and Gertler, M., 1999. Monetary policy and asset price volatility. *Federal Reserve Bank of Kansas Economic Review*, 84(4), p.18-51.

Bernanke, B.S., and Kuttner, K.N., 2005. What explains the stock market's reaction to Federal Reserve policy?. *The Journal of Finance*, 60(3), p.1221-1257.

Bernanke, B.S., and Mihov, I., 1998. Measuring monetary policy. *The Quarterly Journal of Economics*, 113(3), p.869-902.

Bjornland, H.C., and Jacobsen, D.H., 2013. House prices and stock prices: different roles in the US monetary transmission mechanism. *The Scandinavian Journal of Economics*, 115(4), p.1084-1106.

Bjornland, H.C., and Leitemo, K., 2009. Identifying the interdependence between US monetary policy and the stock market. *Journal of Monetary Economics*, 56(2), p.275-282.

Blanchard, O.J., and Quah, D., 1989. The dynamic effects of aggregate demand and supply disturbances. *The American Economic Review*, 79(4), p.655-673.

Bluedorn, J.C., and Bowdler, C., 2011. The open economy consequences of U.S. monetary policy. *Journal of International Money and Finance*, 30(2), p.309-336.

Boivin, J., Kiley, M.T., and Mishkin, F.S., 2010. How has the monetary transmission mechanism evolved over time?. *NBER Working Paper Series*, Working paper 15879.

Bouakez, H., Essid, B., and Normandin, M., 2013. Stock returns and monetary policy: Are there any ties?. *Journal of Macroeconomics*, 36, p.33-50.

Brissimis, S.N., and Magginas, N.S., 2006. Forward-looking information in VAR models and the price puzzle. *Journal of Monetary Economics*, 53(6), p.1225-1234.

Campbell, J.Y., and Vuolteenaho, T., 2004. Bad beta, good beta. *The American Economic Review*, 94(5), 1249-1275.

Canova, F., and Gambetti, L., 2007. Structural changes in the US economy: Is there a role for monetary policy?. *Journal of Economic Dynamics and Control*, 33(2), p.477-490.

Canova, F., and de Nicolo, G., 2002. Monetary disturbances matter for business fluctuations in the G-7. *Journal of Monetary Economics*, 49(6), p.1131-1159.

Carlstrom, C.T., Fuerst, T.S., and Paustian, M., 2009. Monetary policy shocks, Choleski identification, and DNK models. *Journal of Monetary Economics*, 56(7), p.1014-1021.

Castelnuovo, E., 2013. Monetary policy shocks and financial conditions: A Monte Carlo experiment. *Journal of International Money and Finance*, 32(1), p.282-303.

Cecchetti, S.G., Genberg, H., Lipsky, J., and Wadhwani, S., 2000. Asset prices and central bank policy. *International Centre for Monetary and Banking Studies*, Geneva report on the World Economy 2.

Cenesizoglu, T., 2011. Size, book-to-market ratio and macroeconomic news. *Journal of Empirical Finance*, 18(2), p.248-270.

Chang, K., Chen, N., and Leung, C.K.Y., 2011. Monetary policy, term structure and asset return: comparing REIT, housing and stock. *The Journal of Real Estate Finance and Economics*, 43(1), p.221-257.

Chatziantoniou, I., Duffy, D., and Filis, G., 2013. Stock market response to monetary and fiscal policy shocks: Multi-country evidence. *Economic Modelling*, 30, p.754-769.

Cheng, L., and Jin, Y., 2013. Asset prices, monetary policy, and aggregate fluctuations: An empirical investigation. *Economics Letters*, 119(1), p.24-27.

Chordia, T., Huh, S., and Subrahmanyam, A., 2007. The cross-section of expected trading activity. *The Review of Financial Studies*, 20(3), p.709-740.

Christiano, L.J., Eichenbaum, M., and Evans, C., 1996. The effects of monetary policy shocks: Evidence from the flow of funds. *The Review of Economics and Statistics*, 78(1), p.16-34.

Christiano, L.J., Eichenbaum, M., and Evans, C., 1998. Monetary policy shocks: What have we learned and to what end?. *NBER Working Paper Series*, Working paper 6400.

Christiano, L.J., Eichenbaum, M., and Evans, C., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy*, 113(1), p.1-45.

Cook, T., and Hahn, T., 1989. The effect of changes in the federal funds rate target on market interest rates in the 1970s. *Journal of Monetary Economics*, 24(3), p.331-351.

Cooper, R.V.L., 1974. Efficient capital markets and the quantity theory of money. *The Journal of Finance*, 29(3), p.887-908.

Cornell, B., 1983. The money supply announcements puzzle: Review and interpretation. *The American Economic Review*, 73(4), p.644-657.

Crowder, W.J., 2006. The interaction of monetary policy and stock returns. *The Journal of Financial Research*, 29(4), p.523-535.

D'Amico, S., and Farka, M., 2011. The Fed and the stock market: An identification based on intraday futures data. *Journal of Business and Economic Statistics*, 29(1), p.126-137.

Doan, T.A., 2015. RATS handbook for vector autoregressions, 2nd edition, Estima, Evanston, IL.

Doehr, R., and Martinez-Garcia, E., 2015. Monetary policy expectations and economic fluctuations at the zero lower bound. *Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute*, Working paper 240.

Dueker, M.J., 2002. The monetary policy innovation paradox in VARs: A "discrete" explanation. *Federal Reserve Bank of St. Louis Review*, 84(2), p.43-50.

Ehrmann, M., and Fratzscher, M., 2004. Taking stock: Monetary policy transmission to equity markets. *Journal of Money, Credit and Banking*, 36(4), p.719-737.

Enders, W., 2015. Applied econometric time series, 4th edition, John Wiley and Sons, Hoboken, NJ.

Ewing, B.T., Forbes, S.M., and Payne, J.E., 2003. The effects of macroeconomic shocks on sector-specific returns. *Applied Economics*, 35(2), p.201-207.

Fama, E.F., and French, K.R., 1992. The cross-section of expected stock returns. *The Journal of Finance*, 47(2), p.427-465.

Fama, E.F., and French, K.R., 1993. Common risk factors in the returns on stocks and bonds. *Journal of Financial Economics*, 33(1), p.3-56.

Fama, E.F., and French, K.R., 1995. Size and book-to-market factors in earnings and returns. *The Journal of Finance*, 50(1), p.131-155.

Fama, E.F., and French, K.R., 1996. Multifactor explanations of asset pricing anomalies. *The Journal of Finance*, 51(1), p.55-84.

Fama, E.F., and French, K.R., 2015. A five-factor asset pricing model. *Journal of Financial Economics*, 116(1), p.1-22.

Fama, E.F., and French, K.R., 2016. Dissecting anomalies with a five-factor model. *The Review of Financial Studies*, 29(1), p.69-103.

Faust, J., 1998. The robustness of identified VAR conclusions about money. *Carnegie-Rochester Conference Series on Public Policy*, 49(1), p.207-244.

Faust, J., and Leeper, E.M., 1997. When do long-run identifying restrictions give reliable results?. *Journal of Business and Economic statistics*, 15(3), p.345-353.

Fawley, B.W., and Neely, C.J., 2014. The evolution of Federal Reserve policy and the impact of monetary policy surprises on asset prices. *Federal Reserve Bank of St. Louis Review*, 96(1), p.73-109.

Florackis, C., Kontonikas, A., and Kostakis, A., 2014. Stock market liquidity and macro-liquidity shocks: Evidence from the 2007-2009 financial crisis. *Journal of International Money and Finance*, 44, p.97-117.

Fry, R., and Pagan, A., 2011. Sign restrictions in structural vector autoregressions: A critical review. *Journal of Economics Literature*, 49(4), p.938-960.

Furlanetto, F., 2011. Does monetary policy react to asset prices? Some international evidence. *International Journal of Central Banking*, 7(3), p.91-111.

Gali, J., 1992. How well does the IS-LM model fit postwar U.S. data?. *The Quarterly Journal of Economics*, 107(2), p.709-738.

Gali, J., 2014. Monetary policy and rational asset price bubbles. *The American Economic Review*, 104(3), p.721-752.

Gali, J., and Gambetti, L., 2014. The effects of monetary policy on stock market bubbles: Some evidence. *NBER Working Paper Series*, Working paper 19981.

Gambacorta, L., Hofmann, B., and Peersman, G., 2014. The effectiveness of unconventional monetary policy at the zero lower bound: A cross-country analysis. *Journal of Money, Credit and Banking*, 46(4), p.615-642.

Gertler, M. and Gilchrist, S., 1994. Monetary policy, business cycles, and the behaviour of small manufacturing firms. *The Quarterly Journal of Economics*, 109(2), p.309-340.

Giordani, P., 2004. An alternative explanation of the price puzzle. *Journal of Monetary Economics*, 51(6), p.1271-1296.

Gordon, D.B., and Leeper, E.M., 1994. The dynamic impacts of monetary policy: An exercise in tentative identification. *Journal of Political Economy*, 102(6), p.1228-1247.

Goyenko, R.Y., and Ukhov, A.D., 2009. Stock and bond market liquidity: A long-run empirical analysis. *Journal of Financial and Quantitative Analysis*, 44(1), p.189-212.

Guo, H., 2004. Stock prices, firm size, and changes in the federal funds rate target. *The Quarterly Review of Economics and Finance*, 44(4), p.487-507.

Gurkaynak, R.S., Sack, B.P., and Swanson, E.T., 2005. Market-based measures of monetary policy expectations. *Journal of Business and Economic Statistics*, 25(2), p.201-212.

Hahn, J., and Lee, H., 2006. Yield spreads as alternative risk factors for size and book-to-market. *The Journal of Financial and Quantitative Analysis*, 41(2), p.245-269.

Hahn, T., O'Neill, M., and Reyes, G., 2004. Anomalies: Is it the economy? *The Journal of Financial Research*, 27(2), p.273-287.

Hammoudeh, S., Nguyen, D.K., and Sousa, R.M., 2015. US monetary policy and sectoral commodity prices. *Journal of International Money and Finance*, 57, p.61-85.

Hanson, M.S., 2004. The "price puzzle" reconsidered. *Journal of Monetary Economics*, 51(7), p.1385-1413.

Hardouvelis, G.A., 1987. Macroeconomic information and stock prices. *Journal of Economics and Business*, 39(2), p.131-140.

Homa, K.E., and Jaffee, D.M., 1971. The supply of money and common stock prices. *The Journal of Finance*, 26(5), p.1046-1066.

Horowitz, J.L., Loughran, T., and Savin, N.E., 2000. Three analyses of the firm size premium. *Journal of Empirical Finance*, 7(2), p.143-153.

James, C., and Edmister, R.O., 1983. The relation between common stock returns trading activity and market value. *The Journal of Finance*, 38(4), p.1075-1086.

Jansen, D.W., and Tsai, C., 2010. Monetary policy and stock returns: Financing constraints and asymmetries in bull and bear markets. *Journal of Empirical Finance*, 17(5), p.981-990.

Jegadeesh, N., and Titman, S., 1993. Returns to buying winners and selling losers: Implications for stock market efficiency. *The Journal of Finance*, 48(1), p.65-91.

Jensen, G.R., and Mercer, J.M., 2002. Monetary policy and the cross-section of expected stock returns. *The Journal of Financial Research*, 25(1), p.125-139.

Jensen, G.R., and Moorman, T., 2010. Inter-temporal variation in the illiquidity premium. *Journal of Financial Economics*, 98(2), p.338-358.

Kashyap, A.K., Stein, J.C., and Wilcox, D.W., 1993. Monetary policy and credit conditions: Evidence from the composition of external finance. *The American Economic Review*, 83(1), p.78-98.

Keran, M.W., 1971. Expectations, money, and the stock market. *Federal Reserve Bank of St. Louis Review*, 53(1), p.16-31.

Kilian, L., 2011. Structural vector autoregressions. *CEPR Discussion Paper Series*, Discussion paper 8515.

Kim, S., 2001. International transmission of U.S. monetary policy shocks: Evidence from VAR's. *Journal of Monetary Economics*, 48(2), p.339-372.

Kiyotaki, N., and Moore, J., 1997. Credit cycles. *Journal of Political Economy*, 105(2), p.211-248.

Kontonikas, A., and Kostakis, A., 2013. On monetary policy and stock market anomalies. *Journal of Business Finance and Accounting*, 40(7-8), p.1009-1042.

Krusec, D., 2010. The "price puzzle" in the monetary transmission VARs with long-run restrictions. *Economics Letters*, 106(3), p.147-150.

Kurov, A., 2010. Investor sentiment and the stock market's reaction to monetary policy. *Journal of Banking and Finance*, 34(1), p.139-149.

Kurov, A., 2012. What determines the stock market's reaction to monetary policy statements? *Review of Financial Economics*, 21(4), p.175-187.

Kuttner, K.N., 2001. Monetary policy surprises and interest rates: Evidence from the Fed funds futures market. *Journal of Monetary Economics*, 47(3), p.523-544.

Kuttner, K.N., 2011. Monetary policy and asset price volatility: Should we refill the Bernanke-Gertler prescription?. Working paper (unpublished).

Lakonishok, J., Shleifer, A., and Vishny, R.W., 1994. Contrarian investment, extrapolation, and risk. *The Journal of Finance*, 49(5), p.1541-1578.

Lanne, M., and Lutkepohl, H., 2008. Identifying monetary policy shocks via changes in volatility. *Journal of Money, Credit and Banking*, 40(6), p.1131-1149.

Laopodis, N.T., 2013. Monetary policy and stock market dynamics across monetary regimes. *Journal of International Money and Finance*, 33, p.381-406.

Lastrapes, W.D., 1998. International evidence on equity prices, interest rates and money. *Journal of International Money and Finance*, 17(3), p.377-406.

Lee, B., 1992. Causal relations among stock returns, interest rates, real activity, and inflation. *The Journal of Finance*, 47(4), p.1591-1603.

Leeper, E.M., and Zha, T., 2003. Modest policy interventions. *Journal of Monetary Economics*, 50(8), p.1673-1700.

Leeper, E.M., Sims, C.A., and Zha, T., 1996. What does monetary policy do?. *Brookings Papers on Economic Activity*, 27(2), p.1-78.

Li, Y.D., Iscan, T.B., and Xu, K., 2010. The impact of monetary policy shocks on stock prices: Evidence from Canada and the United States. *Journal of International Money and Finance*, 29(5), p.876-896.

Liew, J., and Vassalou, M., 2000. Can book-to-market, size and momentum be risk factors that predict economic growth?. *Journal of Financial Economics*, 57(2), p.221-245.

Lioui, A., and Maio, P., 2014. Interest rate risk and the cross section of stock returns. *Journal of Financial and Quantitative Analysis*, 49(2), p.483-511.

Longstaff, F.A., 2009. Portfolio claustrophobia: Asset pricing in markets with illiquid assets. *The American Economic Review*, 99(4), p.1119-1144.

Lutkepohl, H., and Netsunajev, A., 2015. Structural vector autoregressions with smooth transition in variances. The interaction between U.S. monetary policy and the stock market. Working paper (unpublished).

Mackowiak, B., 2007. External shocks, U.S. monetary policy and macroeconomic fluctuations in emerging markets. *Journal of Monetary Economics*, 54(8), p.2512-2520.

Maio, P., 2014. Another look at the stock return response to monetary policy actions. *Review of Finance*, 18(1), p.321-371.

Mishkin, F.S., 2001. The transmission mechanism and the role of asset prices in monetary policy. *NBER Working Paper Series*, Working paper 8617.

Nyborg, K.G., and Ostberg, P., 2014. Money and liquidity in financial markets. *Journal of Financial Economics*, 112(1), p.30-52.

Park, K., and Ratti, R.A., 2000. Real activity, inflation, stock returns, and monetary policy. *The Financial Review*, 35(2), p.59-78.

Pearce, D.K., and Roley, V.V., 1983. The reaction of stock prices to unanticipated changes in money: A note. *The Journal of Finance*, 38(4), p.1323-1333.

Pearce, D.K., and Roley, V.V., 1985. Stock prices and economic news. *The Journal of Business*, 58(1), p.49-67.

Peng, L., 2005. Learning with information capacity constraints. *The Journal of Financial and Quantitative Analysis*, 40(2), p.307-329.

Perez-Quiros, G., and Timmermann, A., 2000. Firm size and cyclical variations in stock returns. *The Journal of Finance*, 55(3), p.1229-1262.

Pesaran, H.H., and Shin, Y., 1998. Generalized impulse response analysis in linear multivariate models. *Economics Letters*, 58(1), p.17-29.

Petkova, R., 2006. Do the Fama–French factors proxy for innovations in predictive variables?. *The Journal of Finance*, 61(2), p.581-612.

Primiceri, G.E., 2005. Time varying structural vector autoregressions and monetary policy. *Review of Economic Studies*, 72(3), p.821-852.

Rapach, D.E., 2001. Macro shocks and real stock prices. *Journal of Economics and Business*, 53(1), p.5-26.

Rigobon, R., and Sack, B., 2003. Measuring the reaction of monetary policy to the stock market. *The Quarterly Journal of Economics*, 118(2), p.639-669.

Rigobon, R., and Sack, B., 2004. The impact of monetary policy on asset prices. *Journal of Monetary Economics*, 51(8), p.1553-1575.

Rogalski, R.J., and Vinso, J.D., 1977. Stock returns, money supply and the direction of causality. *The Journal of Finance*, 32(4), p.1017-1030.

Romer, C.D., and Romer, D.H., 2004. A new measure of monetary shocks: derivation and implications. *The American Economic Review*, 94(4), p.1055-1084.

Rosa, C., 2011. Words that shake traders. The stock market's reaction to central bank communication in real time. *Journal of Empirical Finance*, 18(5), p.915-934.

Rudebusch, G.D., 1998. Do measures of monetary policy in a VAR make sense? *International Economic Review*, 39(4), p. 907-931.

Rusnak, M., Havranek, T., and Horvath, R., 2013. How to solve the price puzzle? A meta-analysis. *Journal of Money, Credit and Banking*, 45(1), p.37-70.

Sellin, P., 2001. Monetary policy and the stock market: Theory and empirical evidence. *Journal of Economic Surveys*, 15(4), p.491-541.

Sims, C.A., 1980. Macroeconomics and reality. *Econometrica*, 48(1), p.1-48.

Sims, C.A., 1986. Are forecasting models usable for policy analysis? *Federal Reserve Bank of Minneapolis Quarterly Review*, 10(1), p.2-16.

Sims, C.A., 1992. Interpreting the macroeconomic time series facts. The effects of monetary policy. *European Economic Review*, 36(5), p.975-1000.

Sims, C.A., and Zha, T., 1999. Error bands for impulse responses. *Econometrica*, 67(5), p.1113-1155.

Sims, C.A., and Zha, T., 2006. Does monetary policy create recessions?. *Macroeconomic Dynamics*, 10(2), p.231-272.

Smirlock, M., and Yawitz, J., 1985. Asset returns, discount rate changes, and market efficiency. *The Journal of Finance*, 40(4), p.1141-1158.

Sousa, R.M., 2014. Wealth, asset portfolio, money demand and policy rule. *Bulletin of Economic Research*, 66(1), p.95-111.

Stock, J.H., and Watson, M.W., 2001. Vector autoregressions. *Journal of Economic Perspectives*, 15(4), p.101-115.

Strongin, S., 1995. The identification of monetary policy disturbances. Explaining the liquidity puzzle. *Journal of Monetary Economics*, 35(3), p.463-497.

Svensson, L.E.O., 1997. Inflation forecast targeting: Implementing and monitoring inflation targets. *European Economic Review*, 41(6), p.1111-1146.

Thorbecke, W., 1997. On stock market returns and monetary policy. *The Journal of Finance*, 52(2), p.635-654.

Thorbecke, W., and Alami, T., 1994. The effect of changes in the federal funds rate target on stock prices in the 1970s. *Journal of Economics and Business*, 46(1), p.13-19.

Thornton, D.L., 2014. The identification of the response of interest rates to monetary policy actions using market-based measures of monetary policy shocks. *Oxford Economic Papers*, 66(1), p.67-87.

Tsai, C., 2011. The reaction of stock returns to unexpected increases in the federal funds rate target. *Journal of Economics and Business*, 63(2), p.121-138.

Uhlig, H., 2005. What are the effects of monetary policy on output? Results from an agnostic identification procedure. *Journal of Monetary Economics*, 52(2), p.381-419.

Van Dijk, M.A., 2011. Is size dead? A review of the size effect in equity returns. *Journal of Banking and Finance*, 35(12), p.3263-3274.

Vargas-Silva, C., 2008. Monetary policy and the US housing market: A VAR analysis imposing sign restrictions. *Journal of Macroeconomics*, 30(3), p.977-990.

Vassalou, M., 2003. News related to future GDP growth as a risk factor in equity returns. *Journal of Financial Economics*, 68(1), p.47-73.

Vassalou, M., and Xing, Y., 2004. Default risk in equity returns. *The Journal of Finance*, 59(2), p.831-868.

Walsh, C.E., 2003. *Monetary theory and policy*, 2nd ed. Cambridge: MIT Press.

Waud, R.N., 1970. Public interpretation of Federal Reserve discount rate changes: Evidence on the "announcement effect". *Econometrica*, 38(2), p.231-250.

Chapter 3 – Appendix

Table A3.1: Unit root tests

	Panel A		Panel B		Panel C	
	ADF constant	**ADF constant and trend**	**ADF constant**	**ADF constant and trend**	**ADF constant**	**ADF constant and trend**
$\Delta lead_t^a$	-	-	-3.72[14]***	-3.74[14]**	-6.81[11]***	-6.79[11]***
gap_t	-4.34 [3]***	-4.41 [3]***	-4.44[3]***	-4.42[3]***	-3.99[6]***	-3.95[6]**
$\Delta\pi_t^a$	-7.72[11]***	-7.71[11]***	-4.51[14]***	-4.60[14]***	-5.06[11]***	-5.04[11]***
$\pi_t^{comp,a}$	-3.20 [15]**	-3.30 [15]*	-	-	-	-
Δsp_t	-12.41 [0]***	-12.42 [0]***	-12.28[0]***	-12.30[0]***	-10.36[0]***	-10.40[0]***
i_t	-2.15 [5]	-2.73 [5]	-2.50[3]	-2.37[3]	-2.08[3]	-2.35[3]
fff_t	-	-	-2.71[12]*	-2.88[12]	-2.22 [5]	-2.38[5]

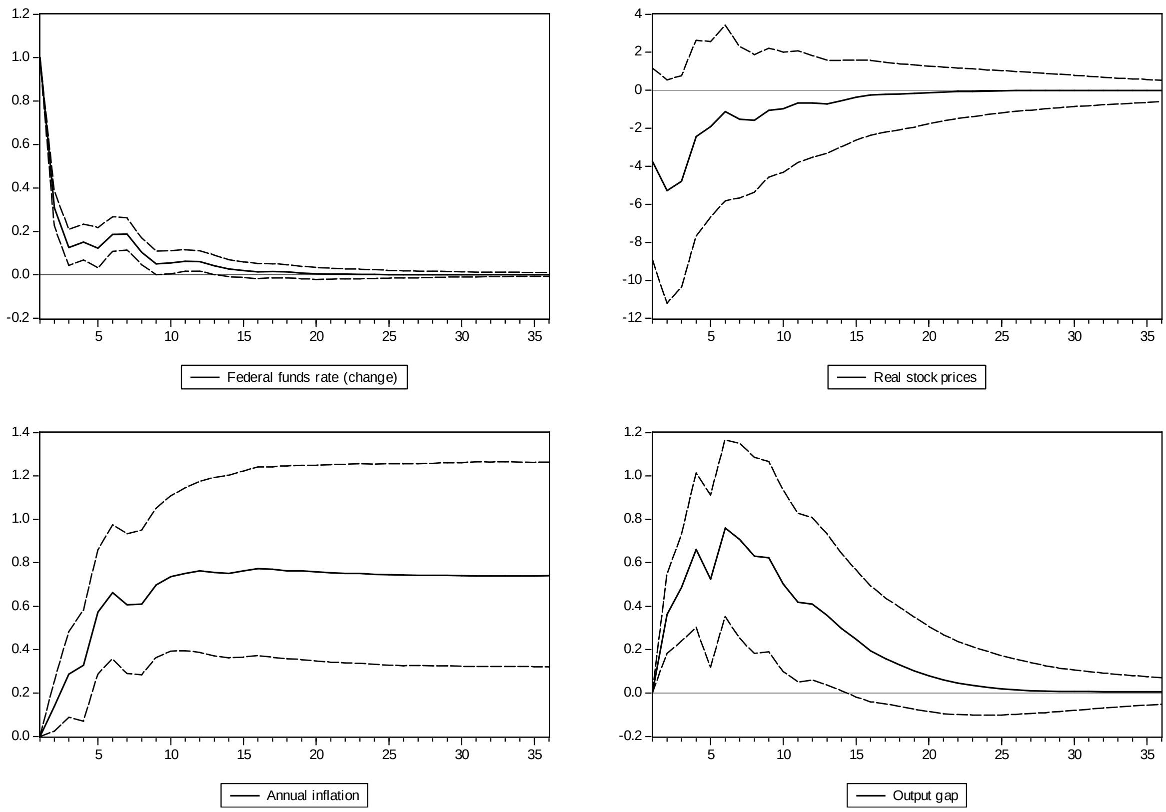
Notes: This table reports the test statistics for the augmented Dickey-Fuller (ADF) unit root tests with (a) constant and (b) constant and trend for the variables in the baseline (Panel A) and augmented (Panel B and Panel C) SVAR models: the first difference of annual growth in the leading economic indicator ($\Delta lead^a_t$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the annual commodity price inflation ($\pi^{comp,a}_t$), the monthly real stock market returns (Δsp_t), the federal funds rate (i_t), and the federal funds futures rate on the 1-month contract (fff_t). The baseline model is estimated over the period 1985:6 - 2007:7 including 6 lags, the unit root tests are reported for the period 1985:1 - 2007:7. The augmented model is estimated over the period 1989:2 - 2007:7 including 6 lags and over the period 1994:2 – 2007:7 including 4 lags, the unit root tests are reported for the period 1988:8 - 2007:7 and 1993:10 – 2007:7. The lag-length of the ADF test, based on the Akaike information criterion, is reported in the brackets. ***, **, * denote 1%, 5% and 10% level of significance, respectively.

Table A3.2: Impulse responses of real stock prices to FFR shock - size and value quintile portfolios

	Panel A: Size-sorted portfolios		Panel B: Value-sorted portfolios	
	Initial period	**Second period**	**Initial period**	**Second period**
1st Quintile	-4.208 (-11.606; 2.895)	-8.909* (-15.757; -3.301)	-11.266* (-18.523; -5.021)	-2.350 (-8.273; 2.464)
2nd Quintile	-9.747* (-16.976; -3.173)	-7.304* (-13.812; -1.746)	-14.394* (-20.115; -9.486)	-2.756 (-7.749; 1.373)
3rd Quintile	-11.956* (-18.632; -5.823)	-5.050 (-11.023; 0.156)	-13.812* (-19.104; -9.160)	-1.956 (-6.642; 2.006)
4th Quintile	-11.126* (-17.174; -5.556)	-2.608 (-8.050; 1.948)	-14.347* (-19.334; -10.219)	-2.450 (-6.714; 1.331)
5th Quintile	-12.655* (-19.165; -7.083)	-2.212 (-7.784; 2.279)	-13.124* (-18.046; -8.533)	-4.252 (-8.987; 0.081)
Spread	4.729 (-0.821; 10.886)	-9.912* (-14.503; -5.132)	-1.327 (-6.039; 3.844)	-0.054 (-3.851; 4.335)

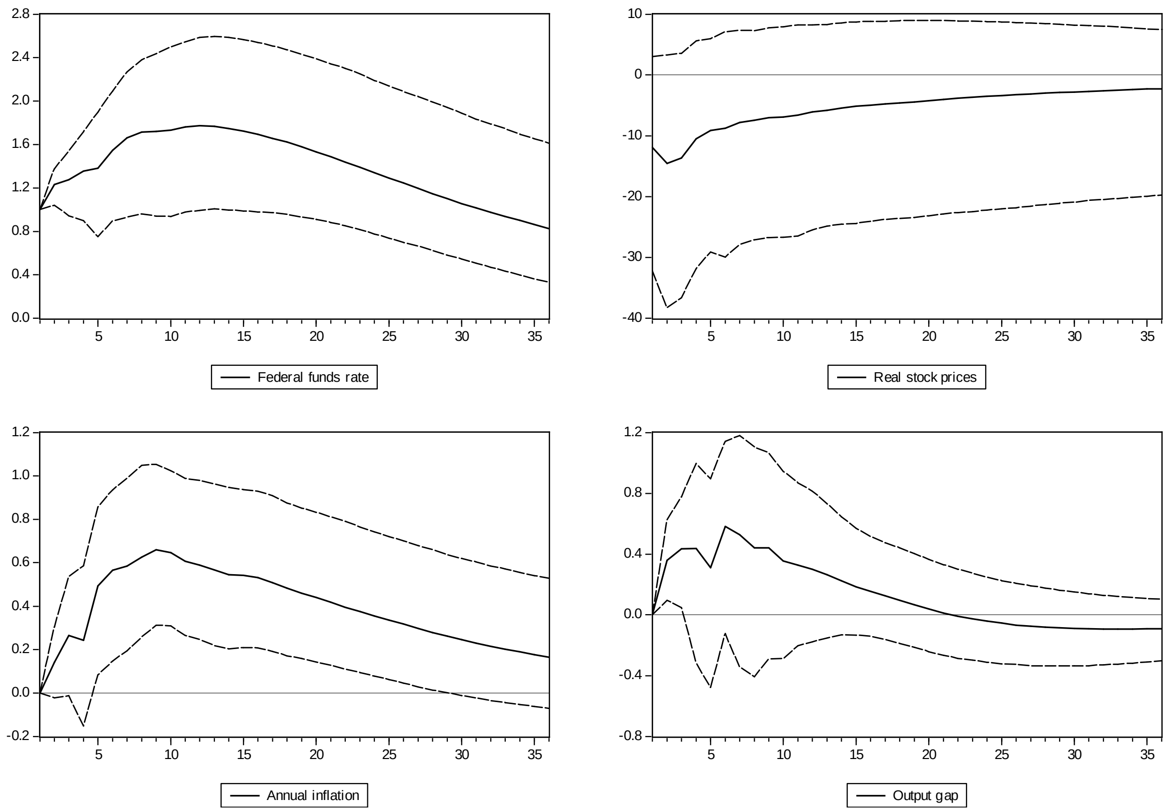
Notes: This table reports the central measures of the initial and second-period impulse responses of real stock prices following a 1-percentage-point contractionary monetary policy shock. Panel A reports the initial (first column) and second-period (second column) impulse responses of stocks across five size-sorted (quintile) portfolios. Panel B reports the initial (first column) and second-period (second column) impulse responses of stocks across five value-sorted (quintile) portfolios. The augmented SVAR model is estimated for each quintile portfolio by replacing real stock market returns in the state vector with relevant real portfolio stock returns over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) in annual change in leading economic indicator ($\Delta lead^{a}_{t-1}$), output gap (gap_t), the first difference in annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the effective federal funds rate (i_t). Dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015) are presented in parentheses. * denote statistically significant impulse responses, i.e. probability bands do not include zero.

Figure A3.1: Impulse responses to FFR shock - change in FFR



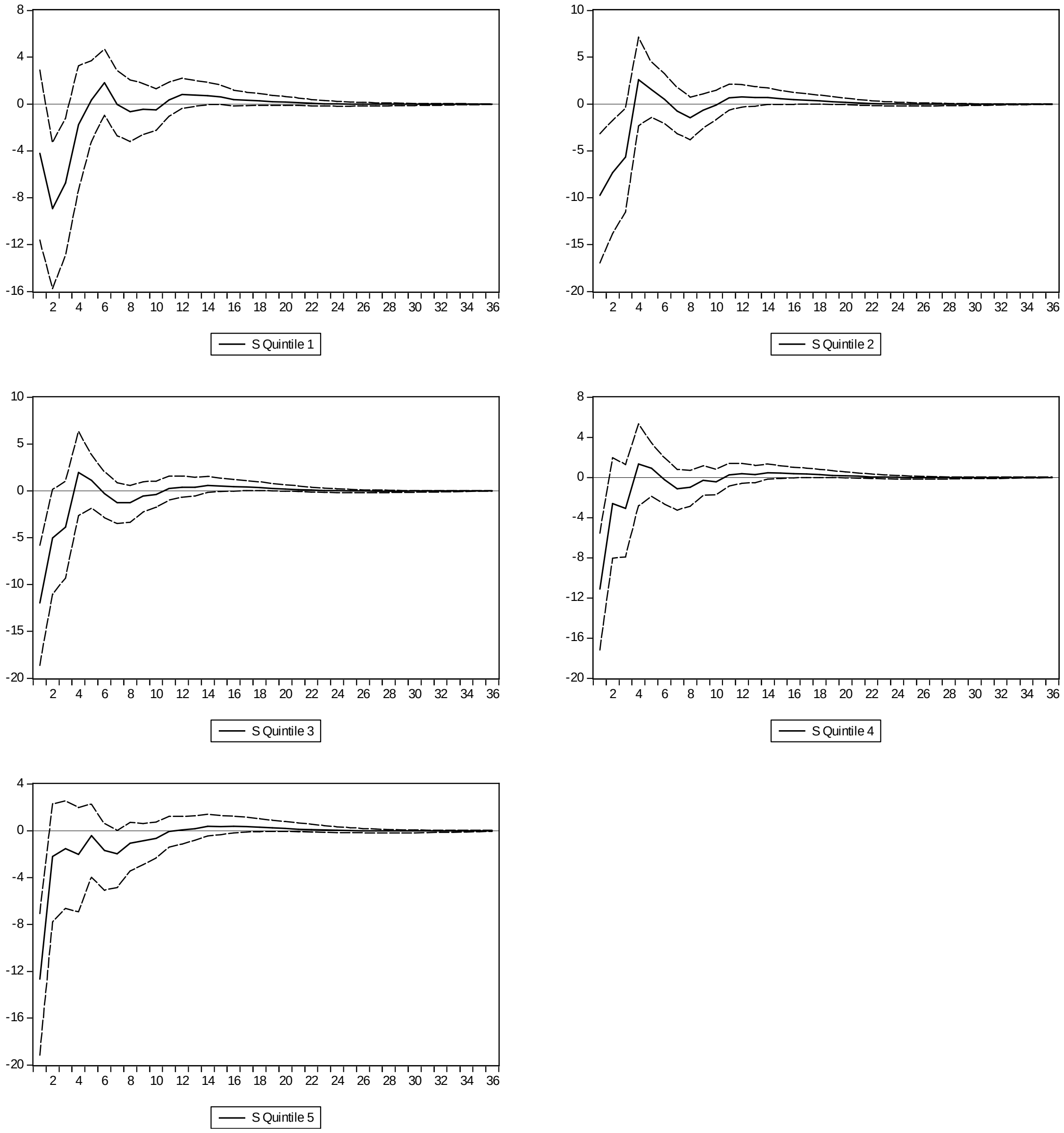
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The baseline SVAR model is estimated over the sample period 1985:6 – 2007:7 including 6 lags. The state vector contains the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a{}_t$), the annual commodity price inflation ($\pi_t^{comp,a}$), monthly real stock market returns (Δsp_t) and the first difference of the federal funds rate (Δi_t). The dummy variables that take value of one in 1987:10 and 2001:9, respectively, and zero otherwise are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

Figure A3.2: Impulse responses to FFR shock - level of annual inflation



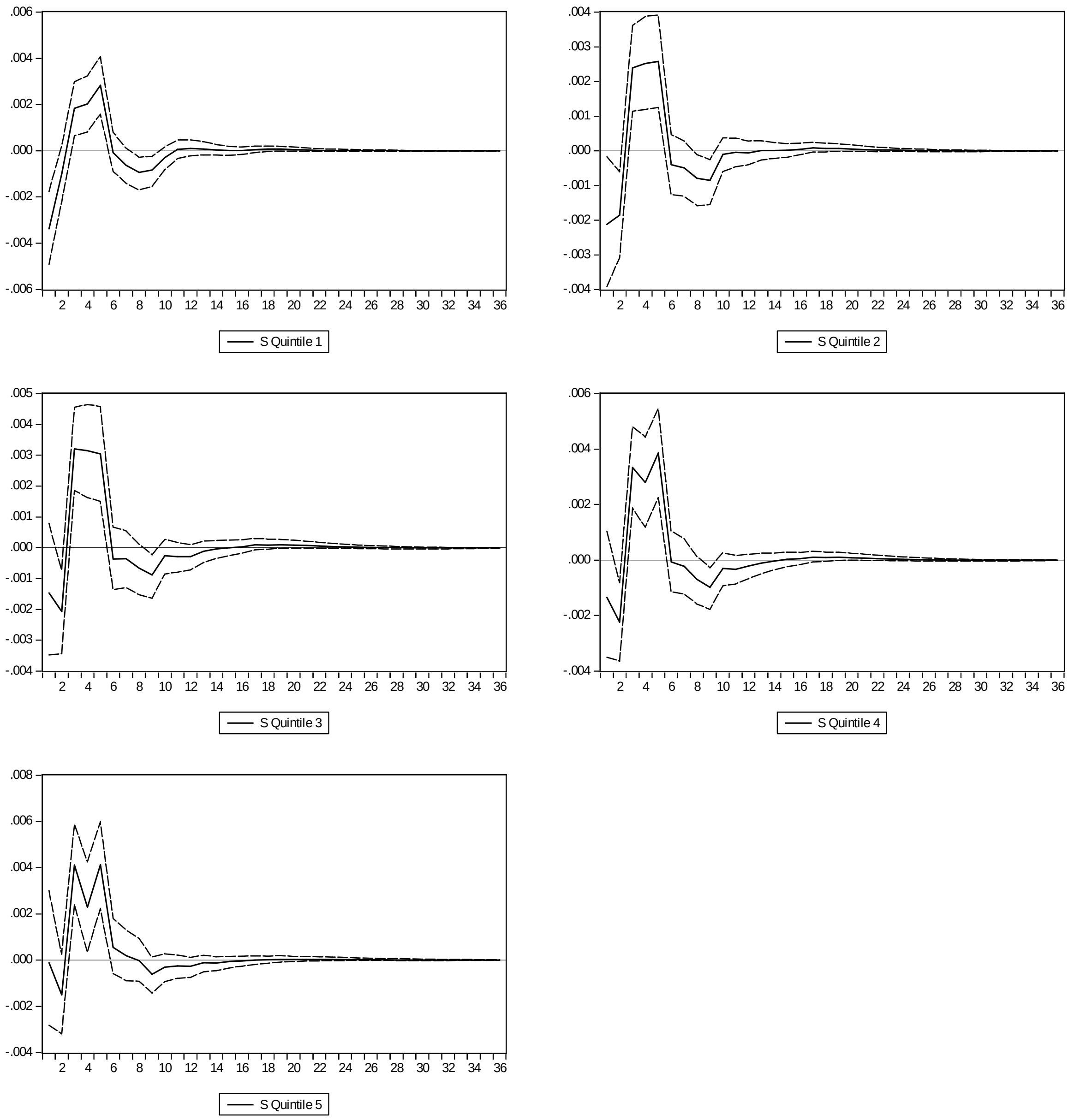
Notes: This figure plots the central measures of impulse responses (median) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The baseline SVAR model is estimated over the sample period 1985:6 – 2007:7 including 6 lags. The state vector contains the output gap (gap_t), the annual inflation ($\pi^a{}_t$), the annual commodity price inflation ($\pi_t^{comp,a}$), monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). See also Figure A3.1 notes.

Figure A3.3: Impulse responses of real stock prices to FFR shock - size quintile portfolios



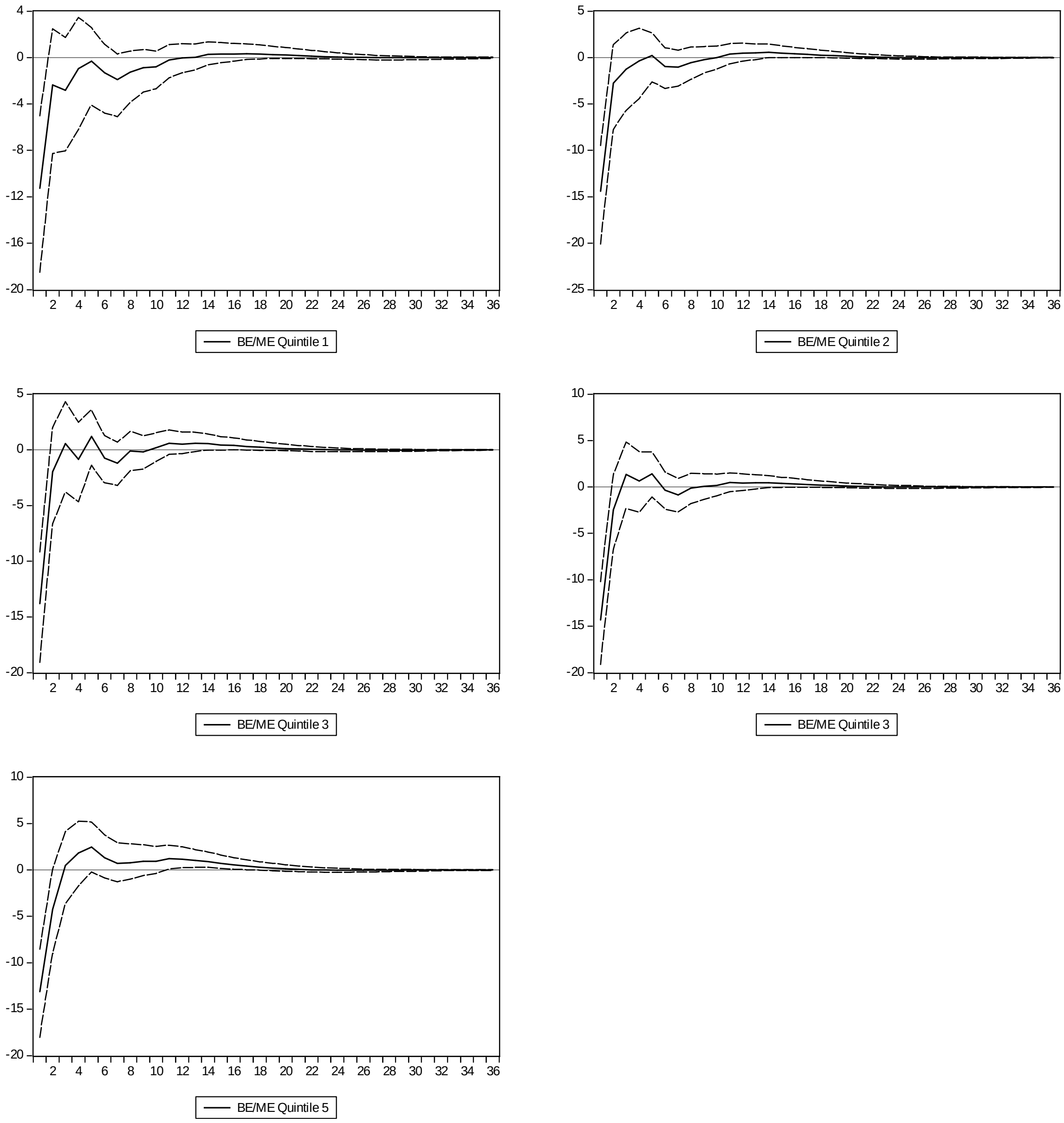
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across five size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

Figure A3.4: Impulse responses of the federal funds rate to SP shock - size quintile portfolios



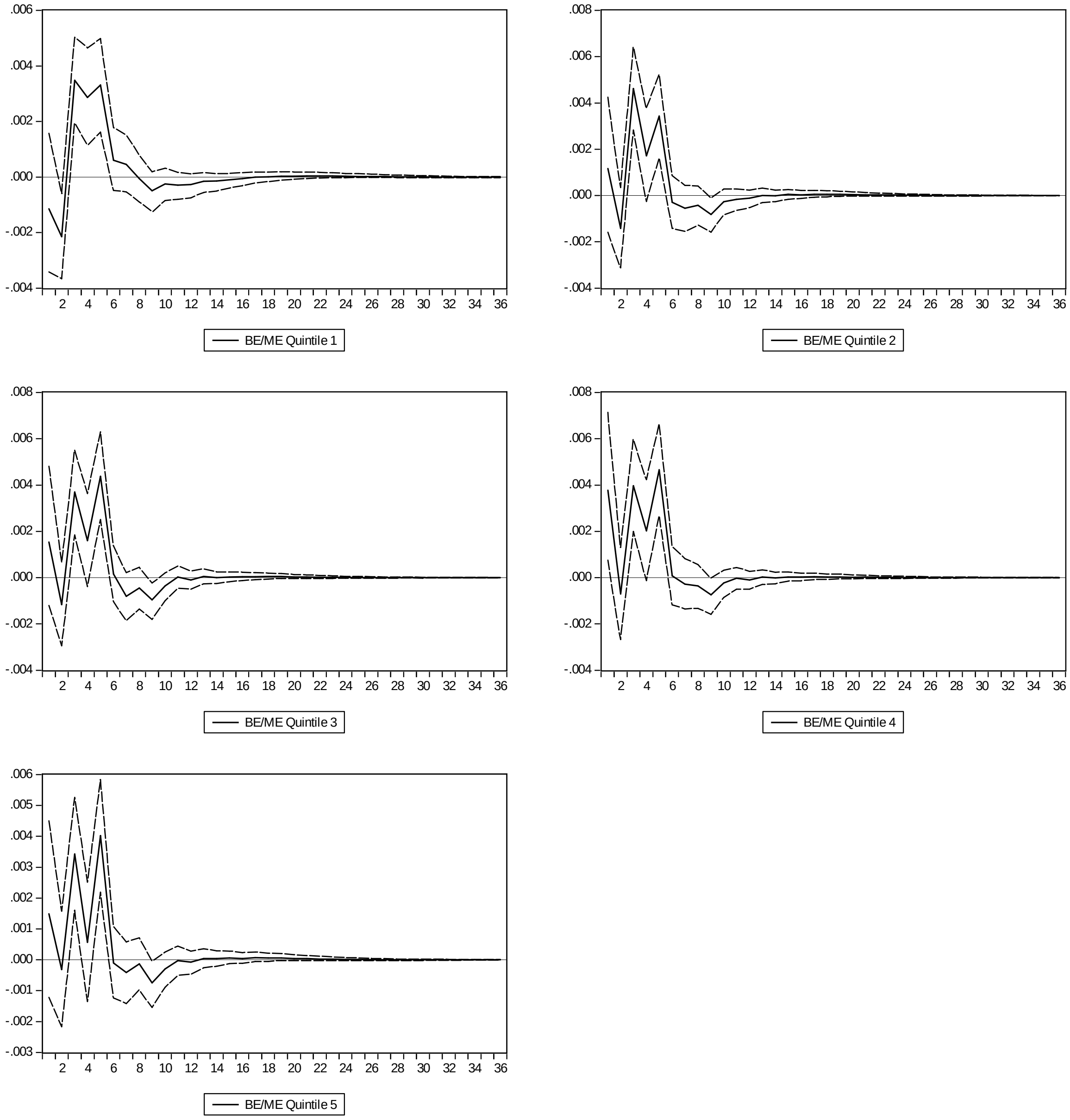
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across five size-sorted portfolios. The augmented SVAR model is estimated for each size quintile portfolio by replacing real stock market returns in the stated vector with the relevant portfolio real stock returns. See also Figure A3.3 notes.

Figure A3.5: Impulse responses of real stock prices to FFR shock - value quintile portfolios



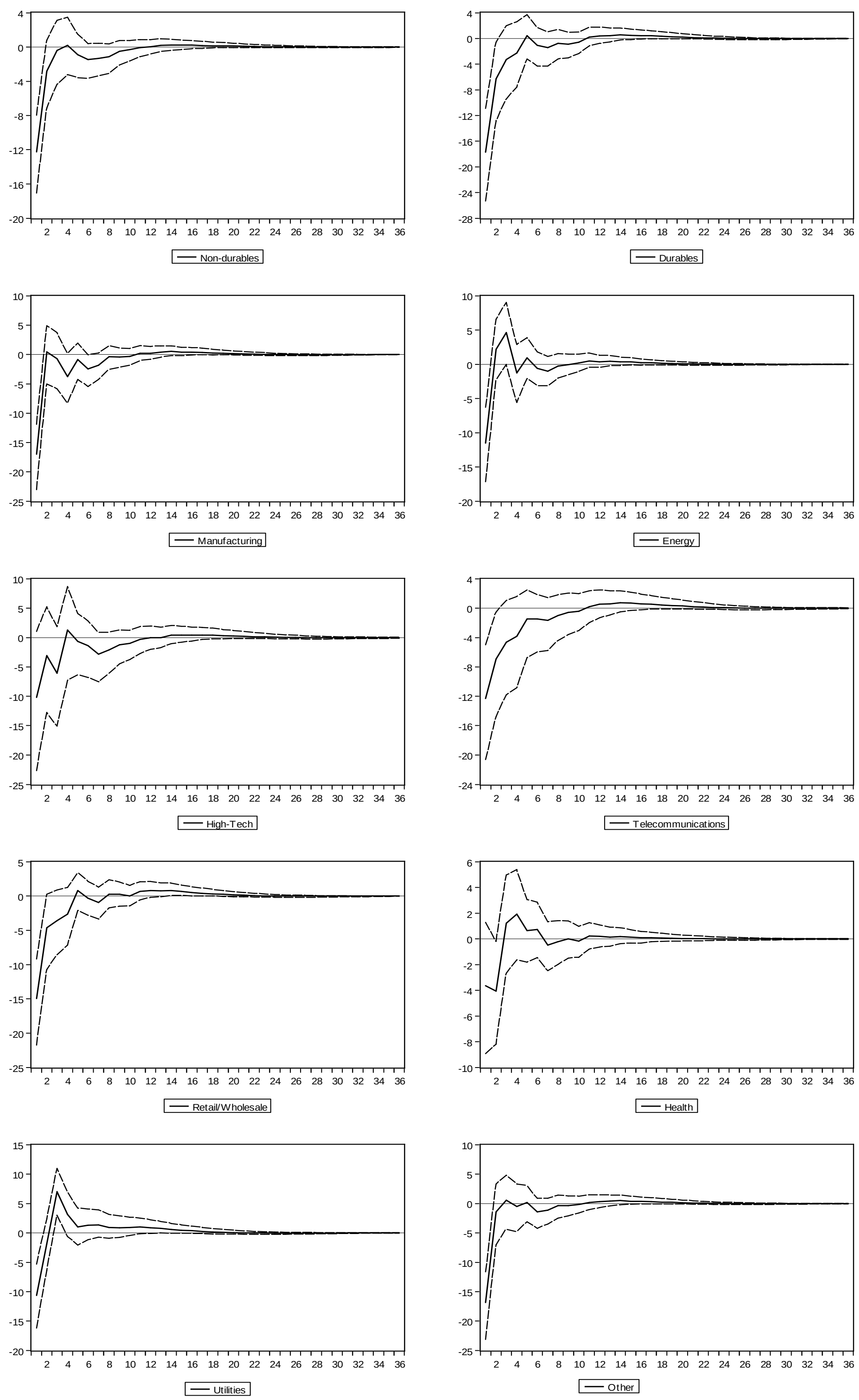
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across five value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.3 notes.

Figure A3.6: Impulse responses of the federal funds rate to SP shock - value quintile portfolios



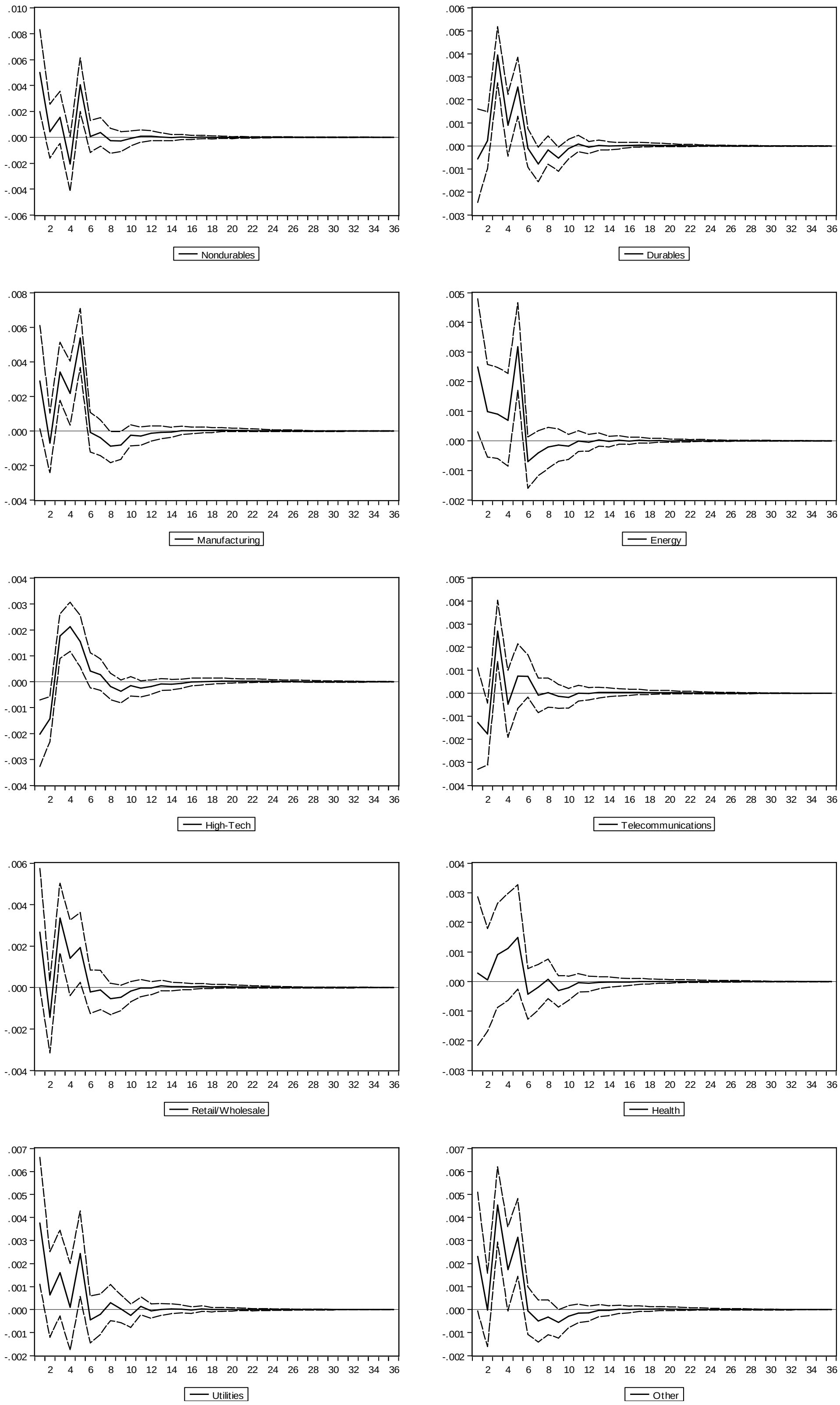
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across ten value-sorted portfolios. The augmented SVAR model is estimated for each value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.3 notes.

Figure A3.7: Impulse responses of real stock prices to FFR shock - industry portfolios



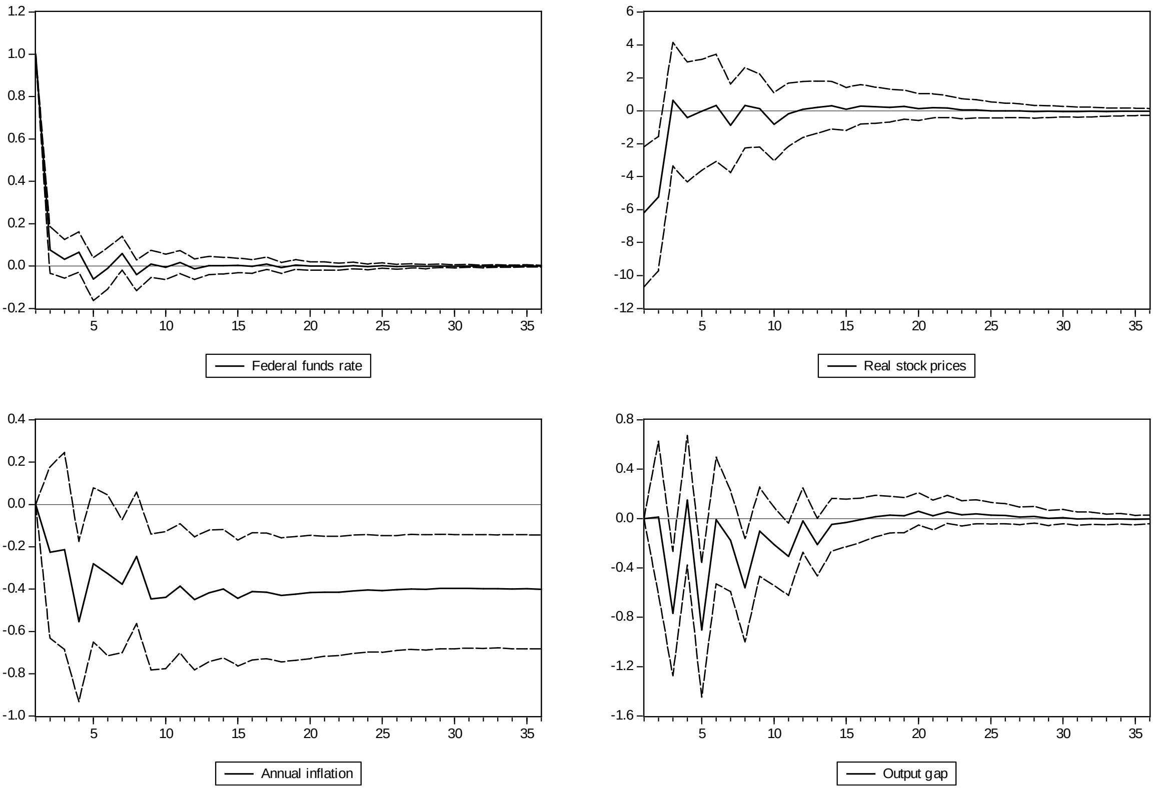
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across ten industry portfolios following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each industry portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.3 notes.

Figure A3.8: Impulse responses of the federal funds rate to SP shock - industry portfolios



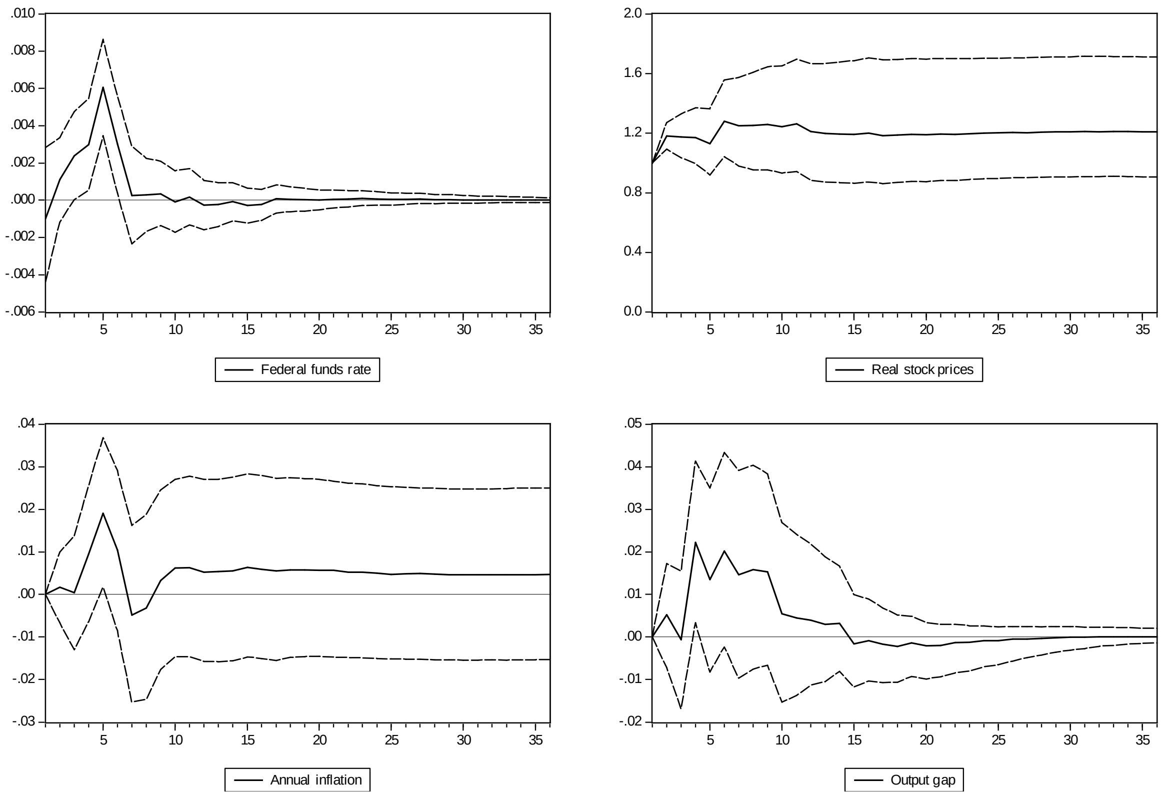
Notes: This figure plots central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across ten industry portfolios. The augmented SVAR model is estimated for each industry portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.3 notes.

Figure A3.9: Impulse responses to FFR shock – VAR(6)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 6 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

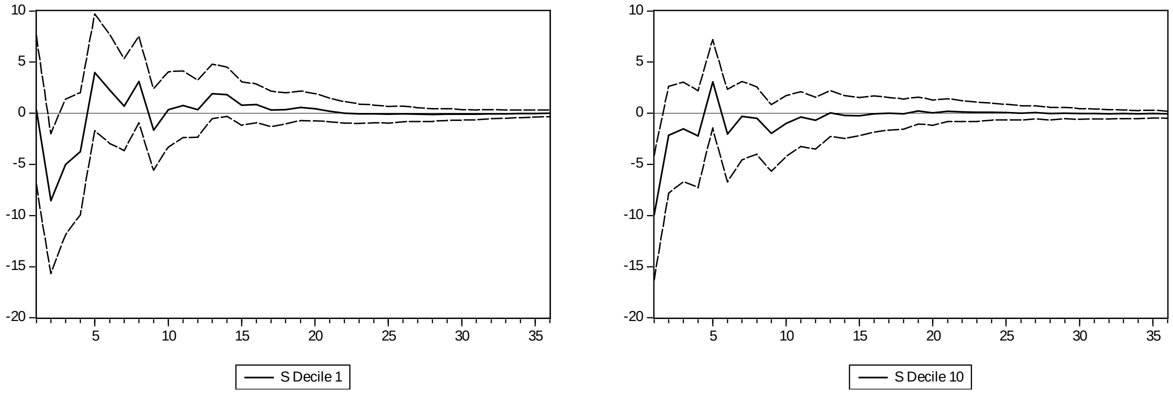
Figure A3.10: Impulse responses to SP shock – VAR(6)



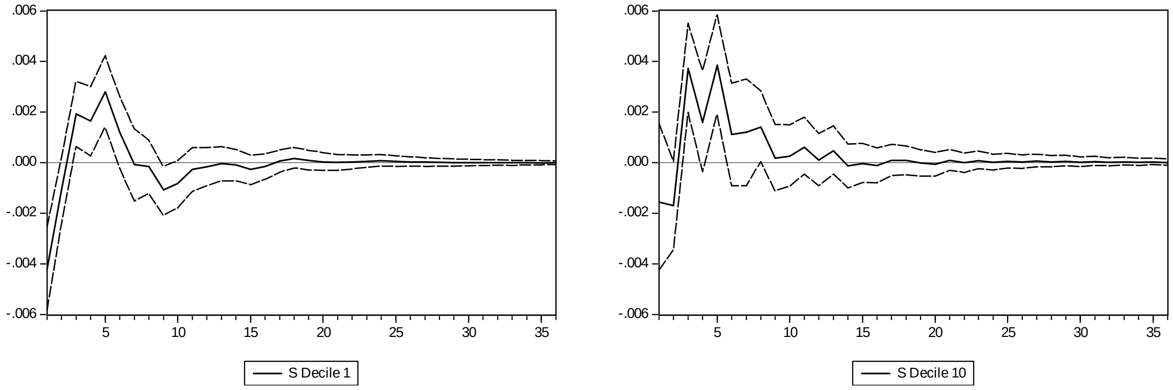
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.9 notes.

Figure A3.11: Impulse responses to FFR and SP shocks – VAR(6) (size decile portfolios)

Panel A: FFR shock



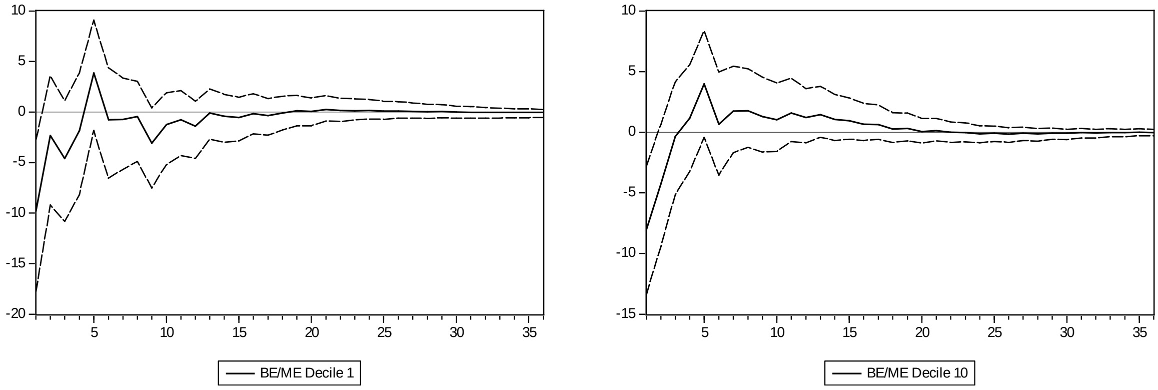
Panel B: SP shock



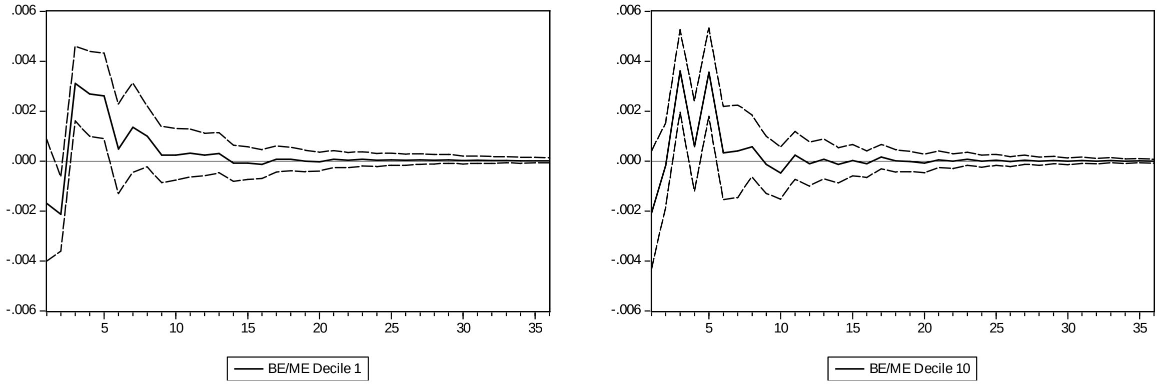
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.9 notes.

Figure A3.12: Impulse responses to FFR and SP shocks – VAR(6) (value decile portfolios)

Panel A: FFR shock

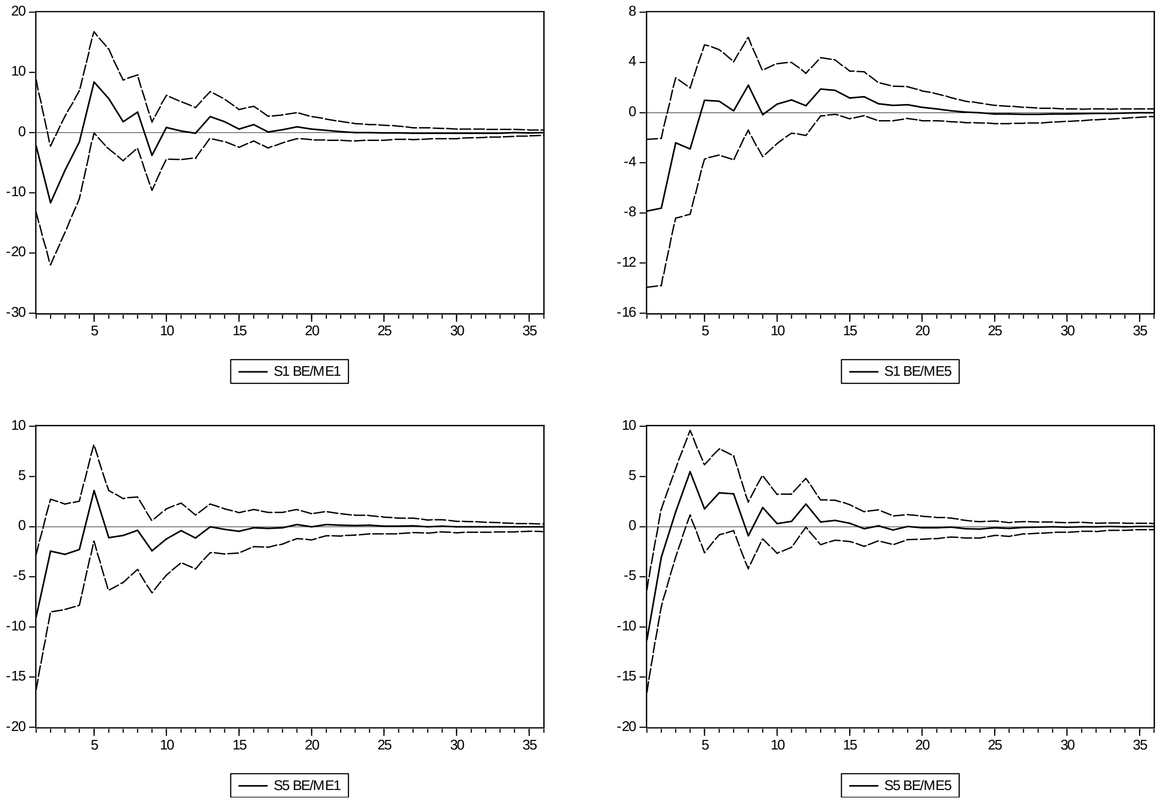


Panel B: SP shock



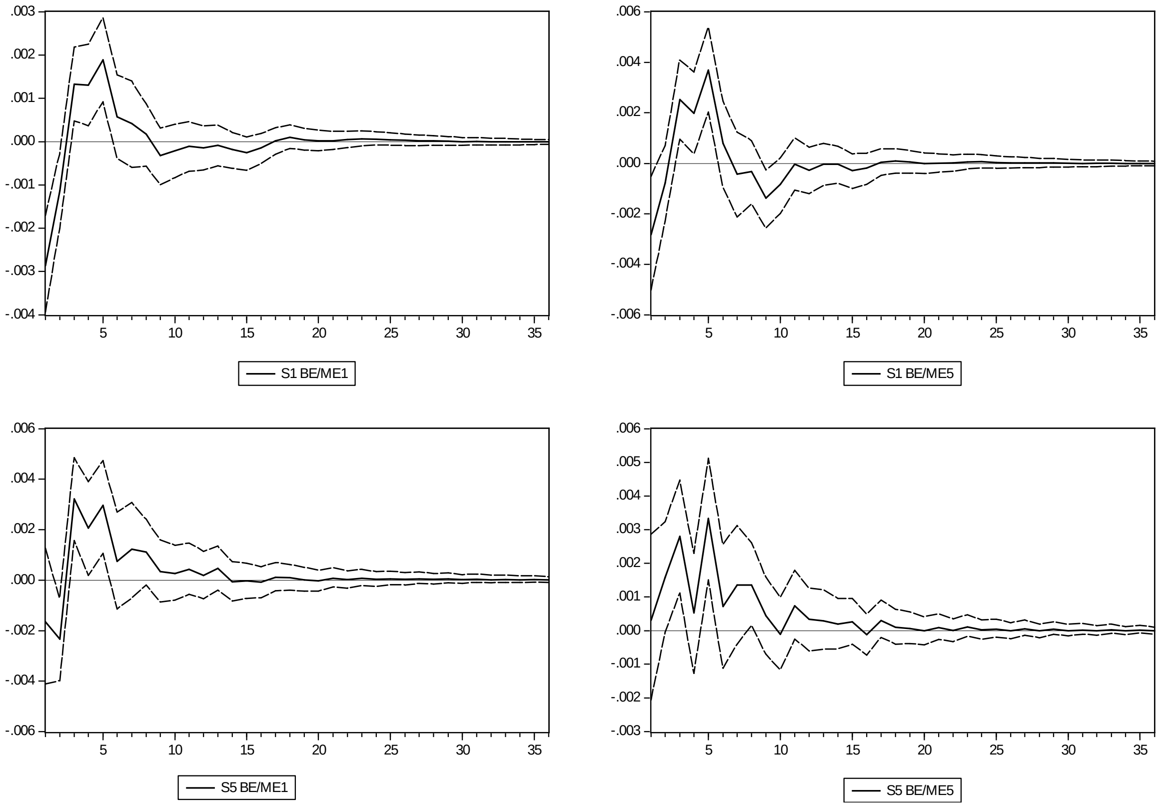
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.9 notes.

Figure A3.13: Impulse responses of real stock prices to FFR shock – VAR(6) (size-value quintile portfolios)



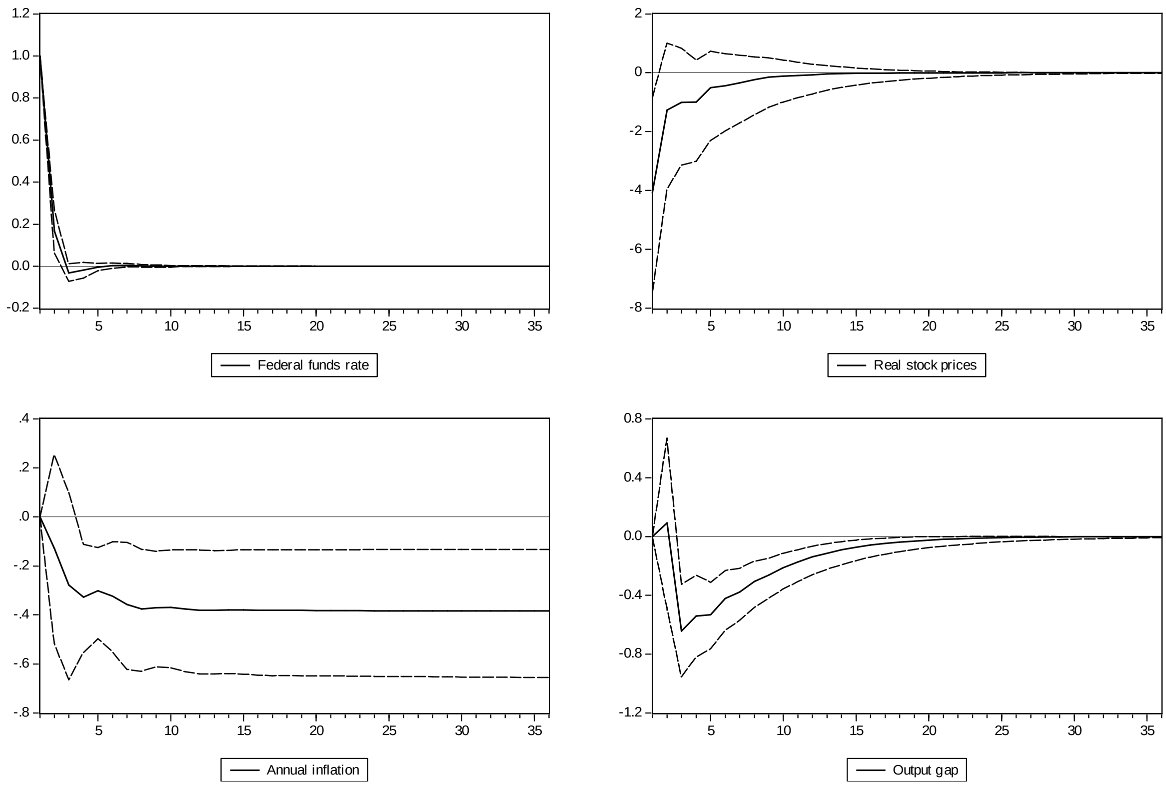
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.9 notes.

Figure A3.14: Impulse responses of the federal funds rate to SP shock – VAR(6) (size-value quintile portfolios)



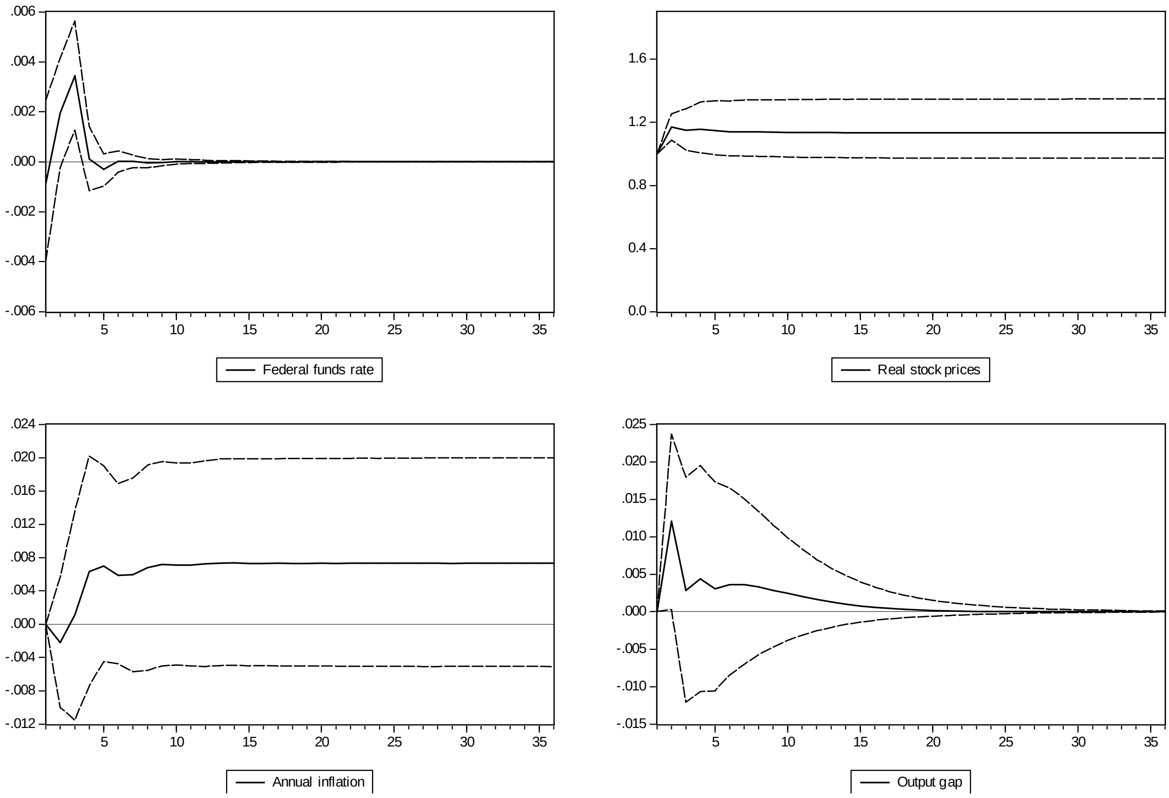
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.9 notes.

Figure A3.15: Impulse responses to FFR shock – VAR(2)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 2 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

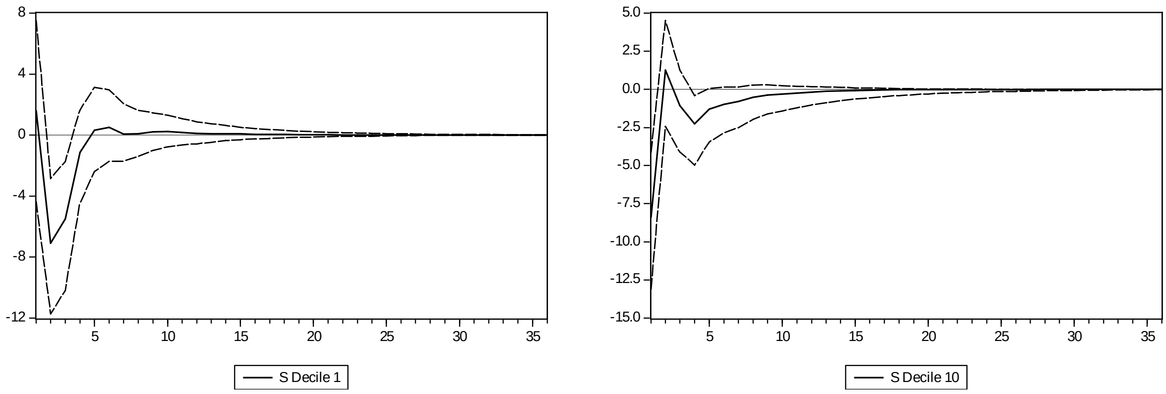
Figure A3.16: Impulse responses to SP shock – VAR(2)



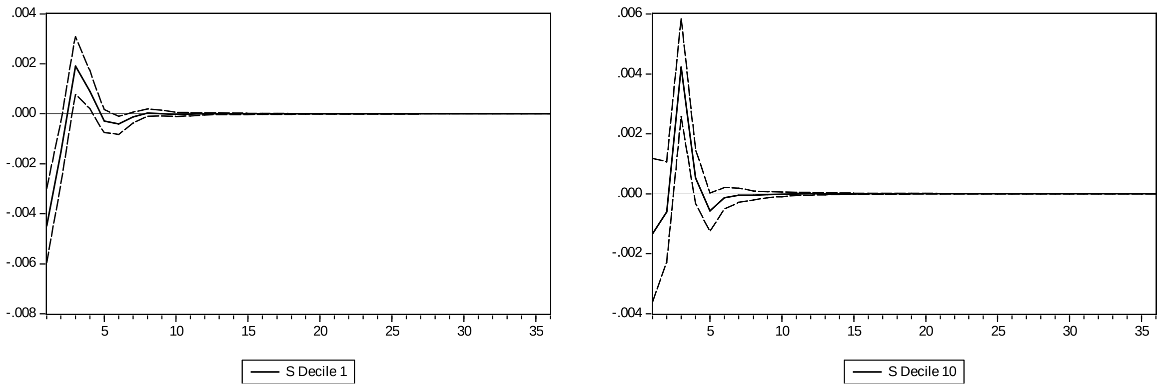
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.15 notes.

Figure A3.17: Impulse responses to FFR and SP shocks - VAR(2) (size decile portfolios)

Panel A: FFR shock



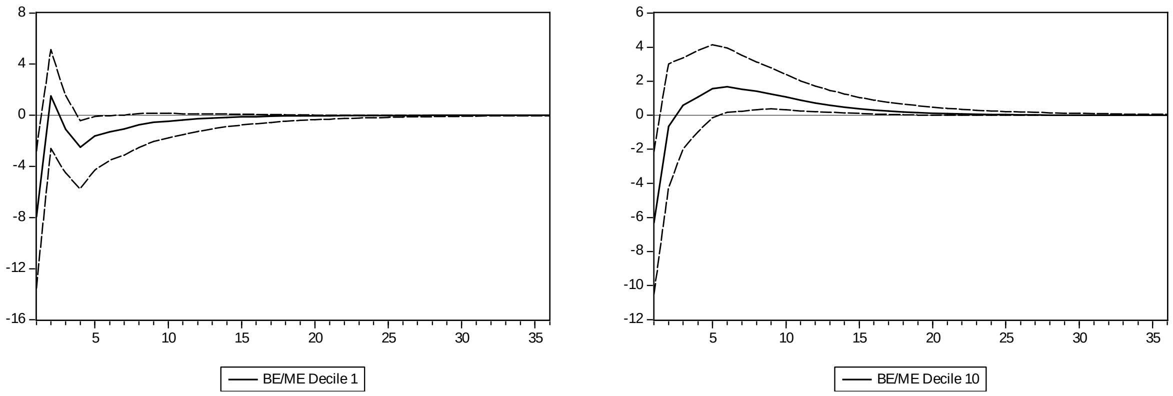
Panel B: SP shock



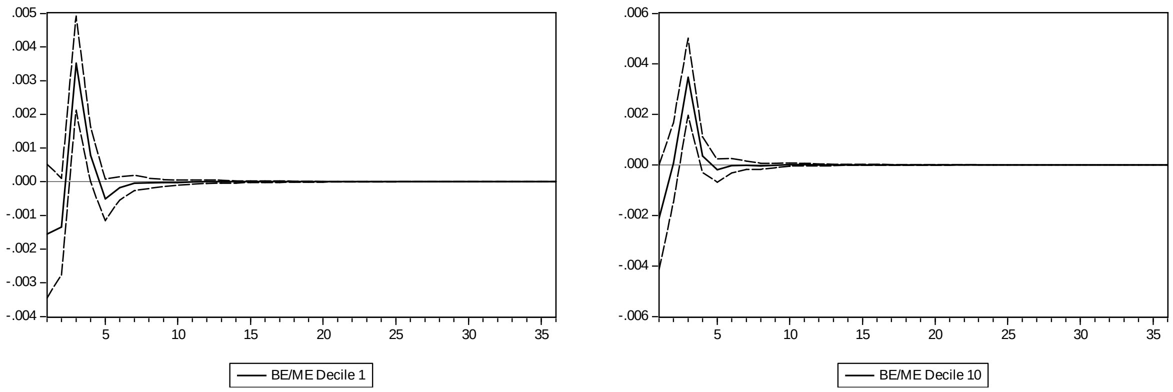
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.15 notes.

Figure A3.18: Impulse responses to FFR and SP shocks - VAR(2) (value decile portfolios)

Panel A: FFR shock

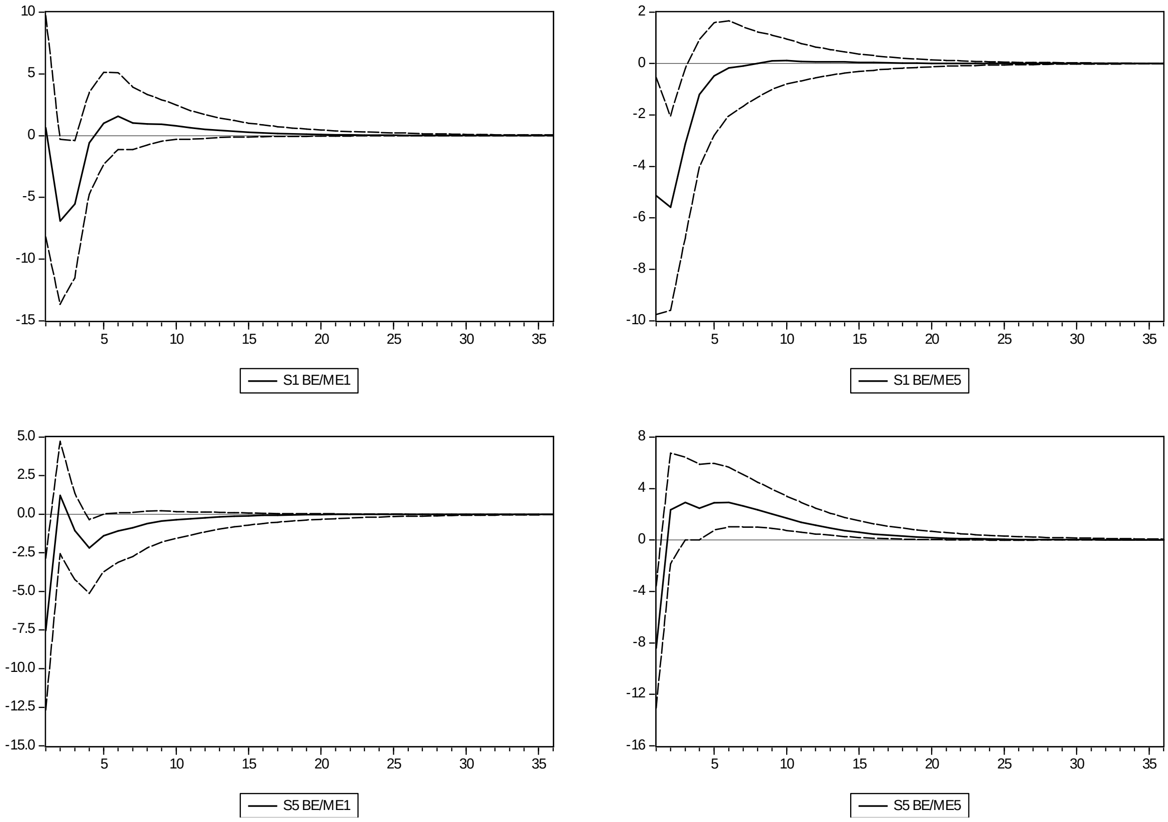


Panel B: SP shock



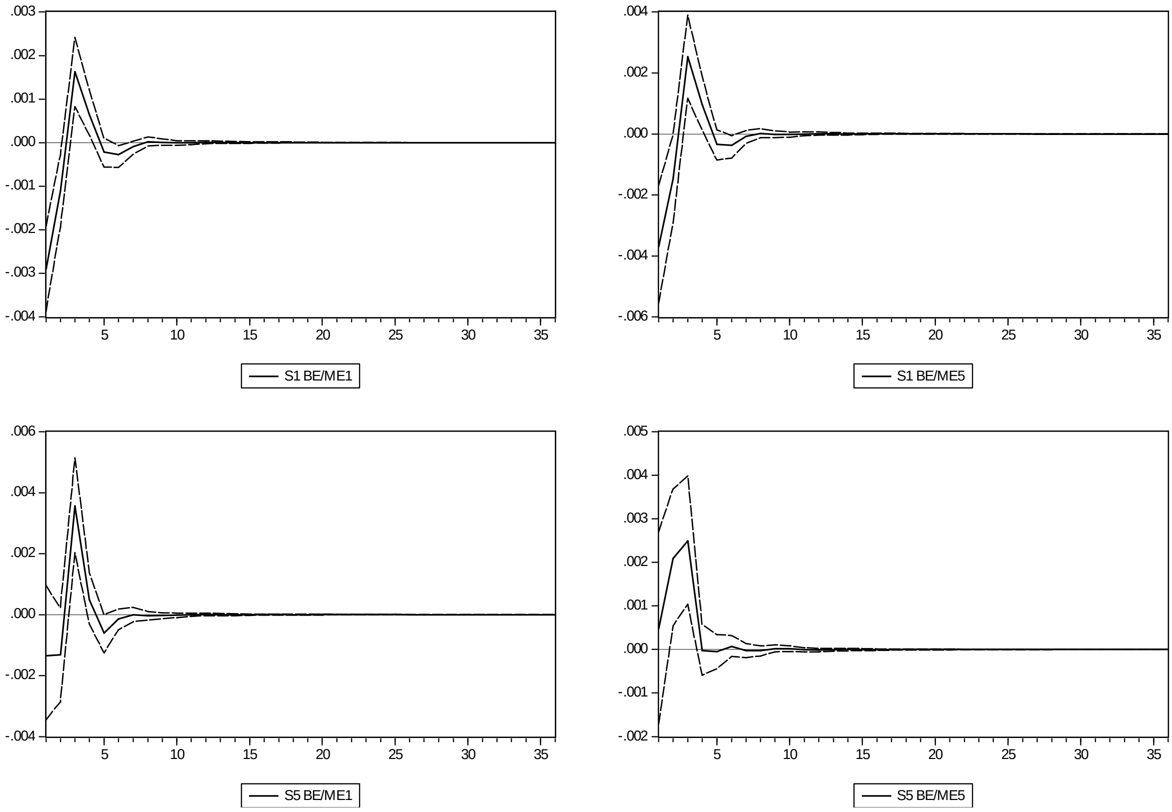
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.15 notes.

Figure A3.19: Impulse responses of real stock prices to FFR shock – VAR(2) (size-value quintile portfolios)



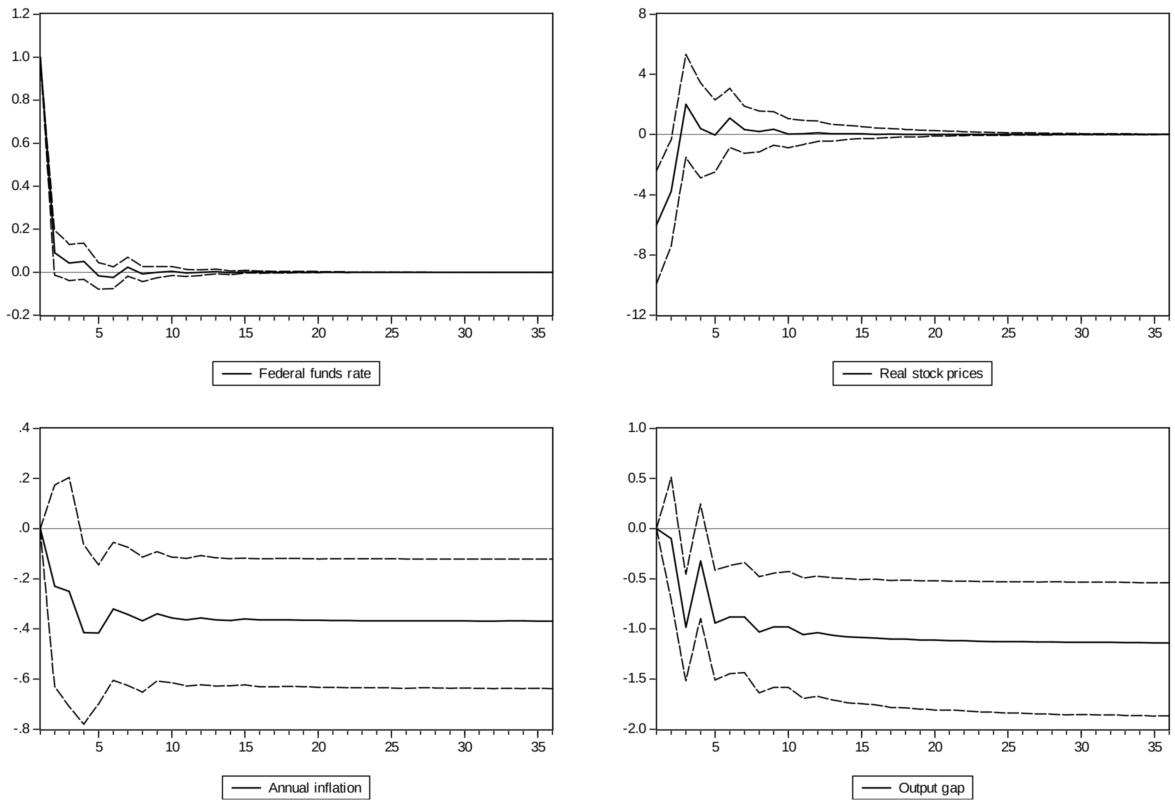
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.15 notes.

Figure A3.20: Impulse responses of the federal funds rate to SP shock – VAR(2) (size-value quintile portfolios)



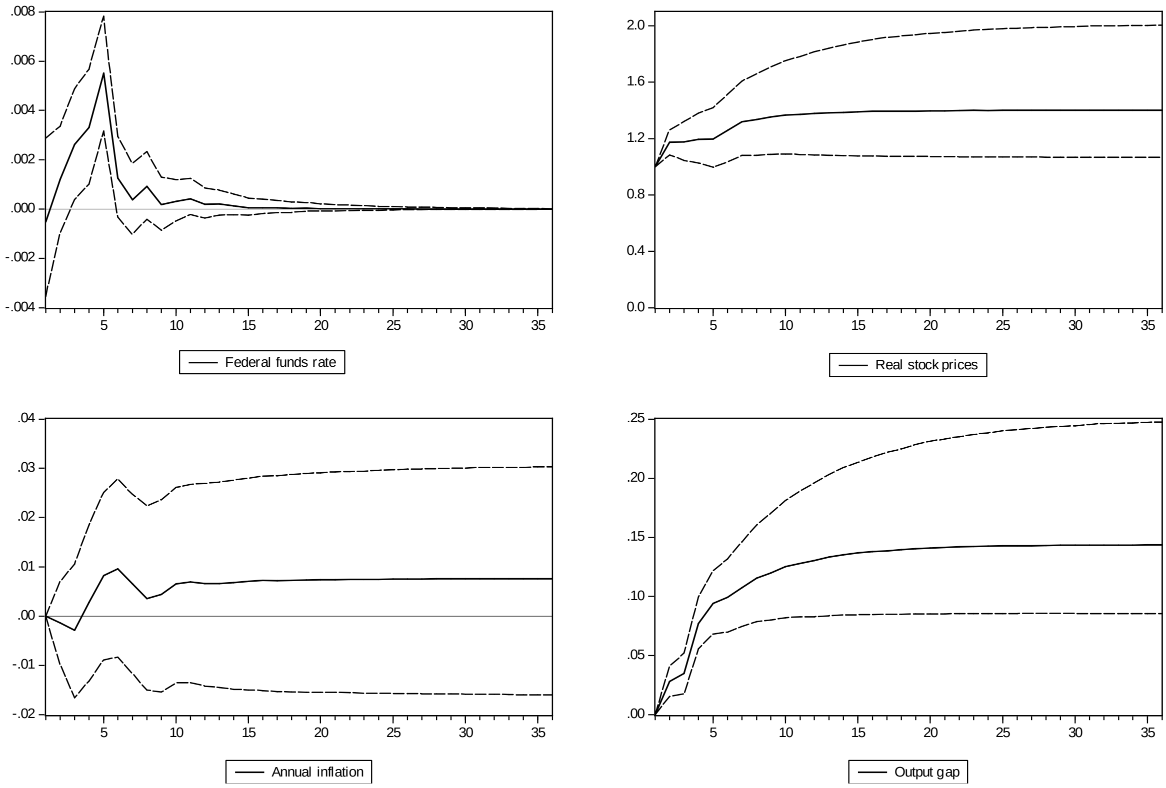
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate (i_t) following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.15 notes.

Figure A3.21: Impulse responses to FFR shock - output gap - quadratic trend



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap calculated using quadratic trend (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

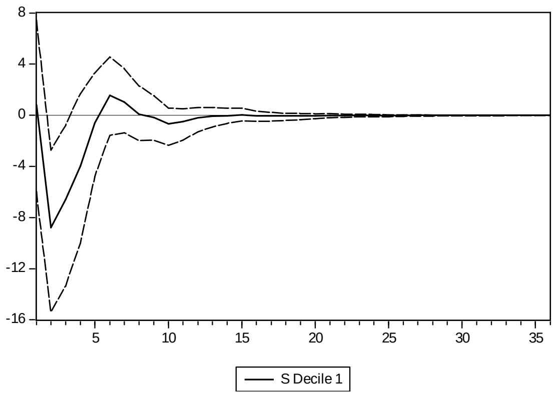
Figure A3.22: Impulse responses to SP shock - output gap - quadratic trend

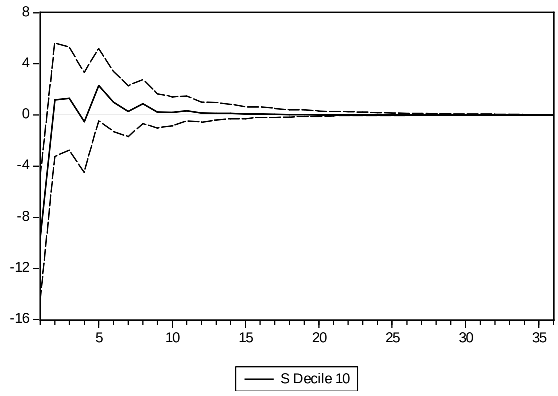


Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.21 notes.

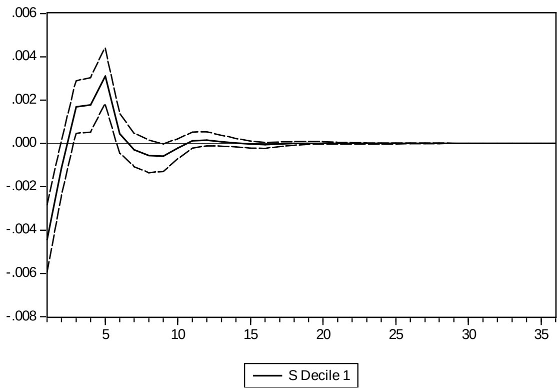
Figure A3.23: Impulse responses to FFR and SP shocks – output gap - quadratic trend (size decile portfolios)

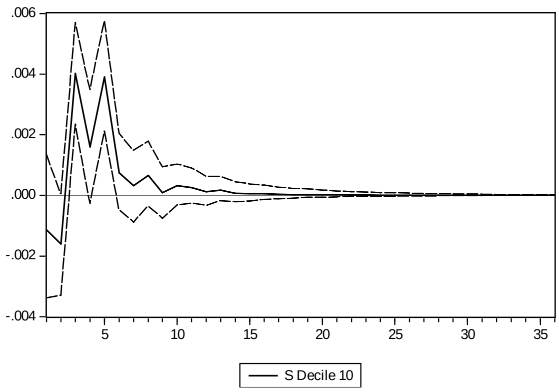
Panel A: FFR shock





Panel B: SP shock

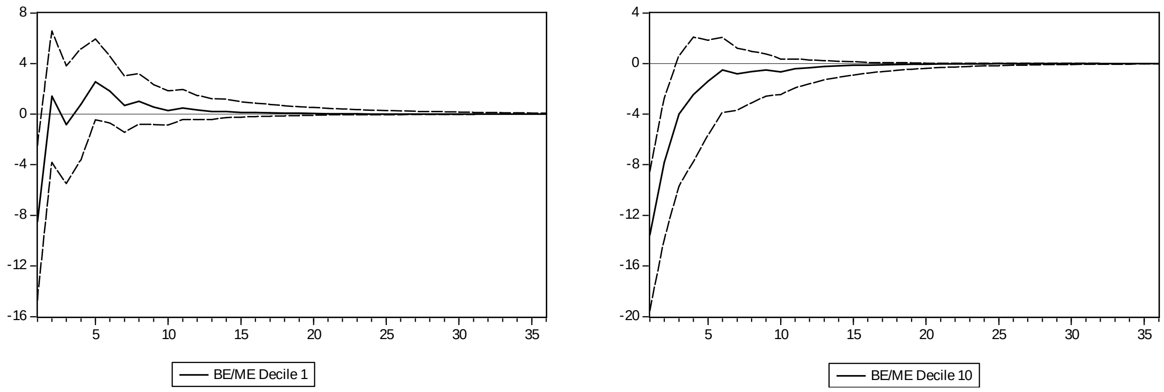




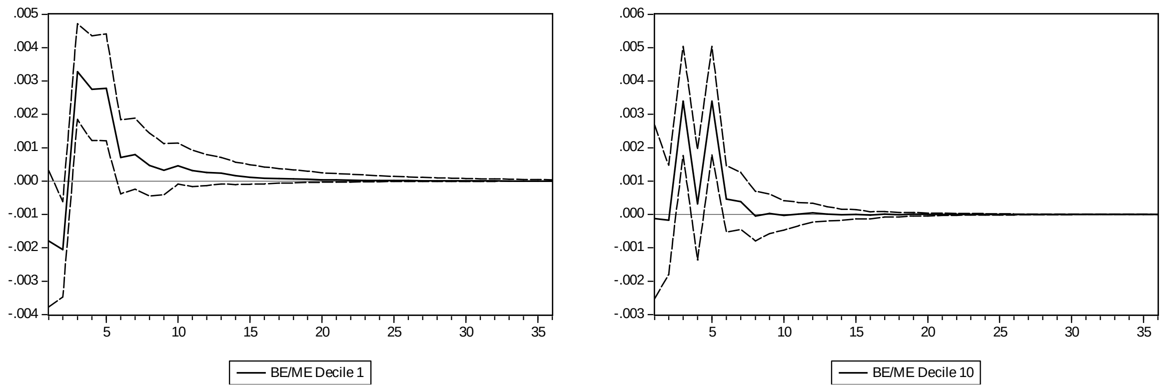
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.21 notes.

Figure A3.24: Impulse responses to FFR and SP shocks - output gap - quadratic trend (value decile portfolios)

Panel A: FFR shock

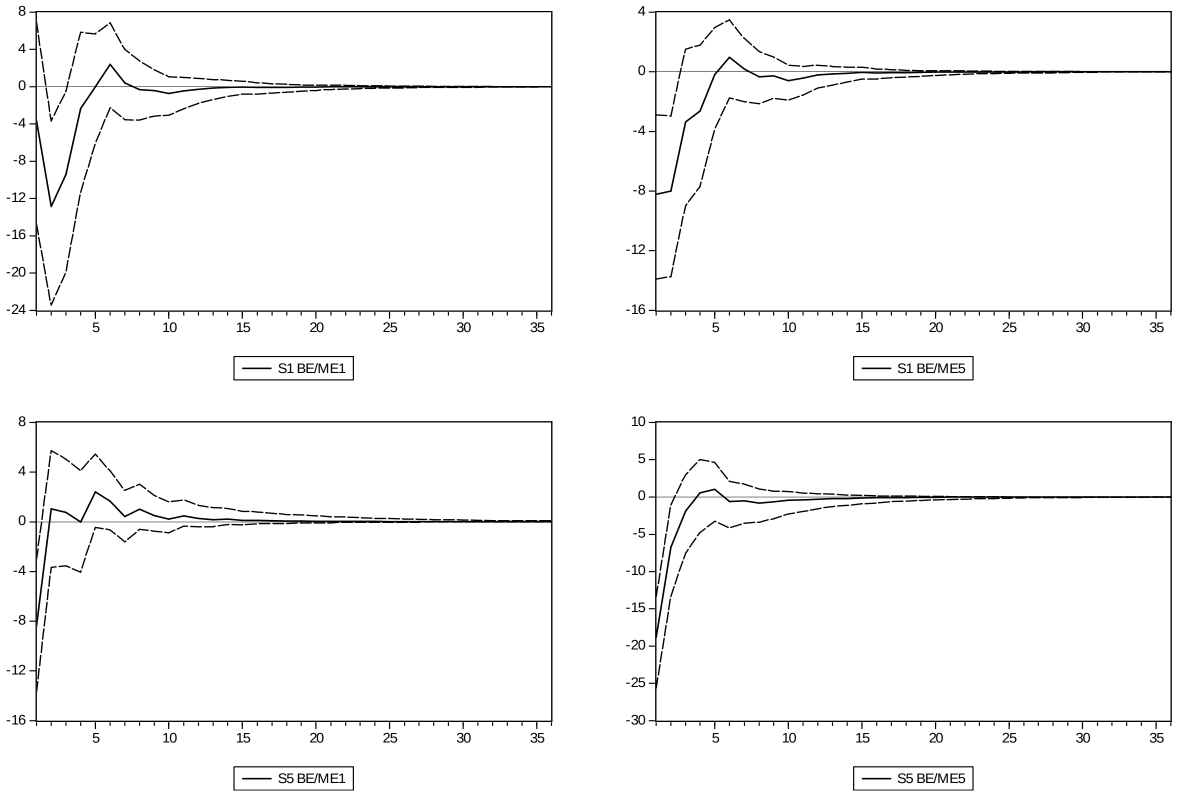


Panel B: SP shock



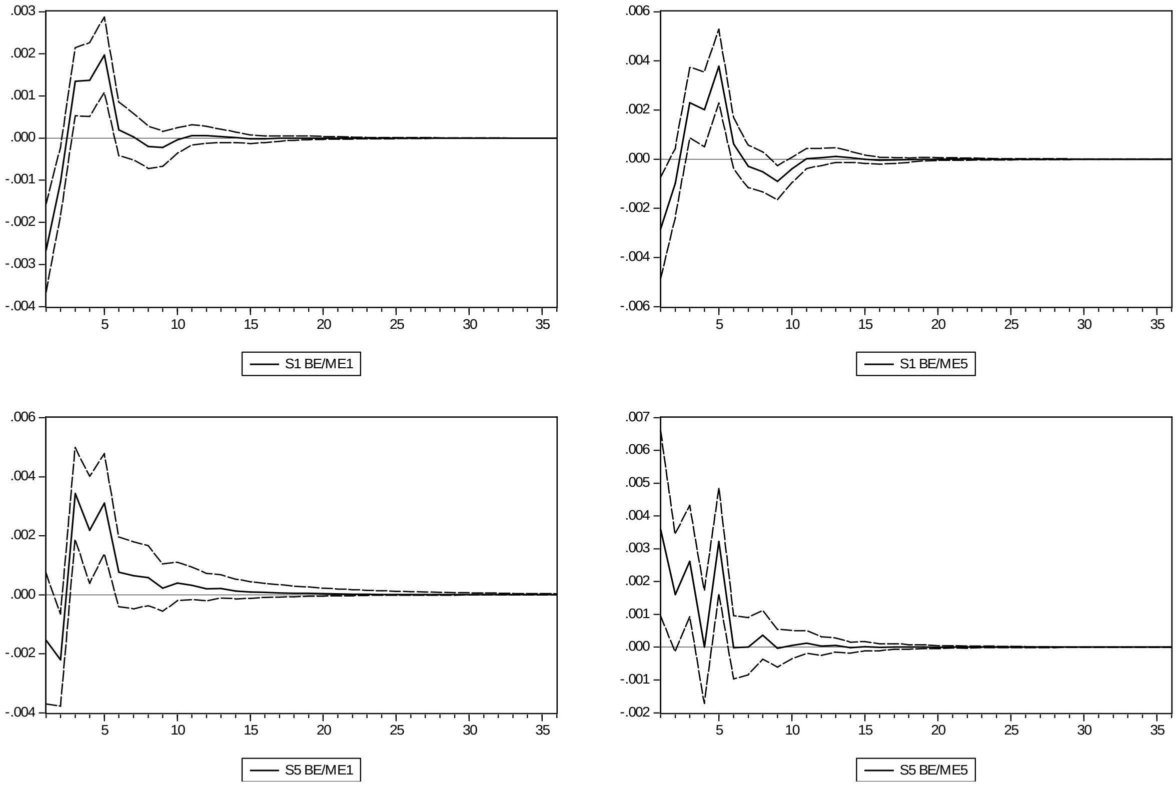
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.21 notes.

Figure A3.25: Impulse responses of real stock prices to FFR shock - output gap – quadratic trend (size-value quintile portfolios)



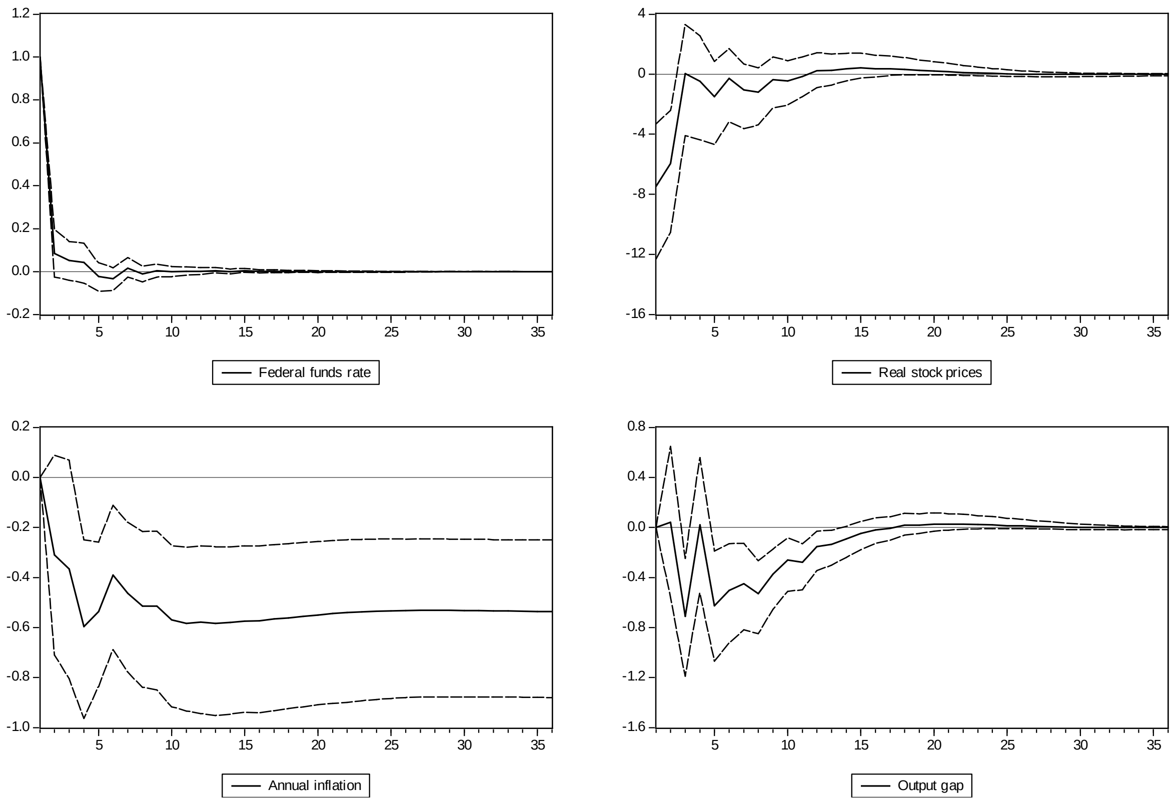
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.21 notes.

Figure A3.26: Impulse responses of the federal funds rate to SP shock - output gap – quadratic trend (size-value quintile portfolios)



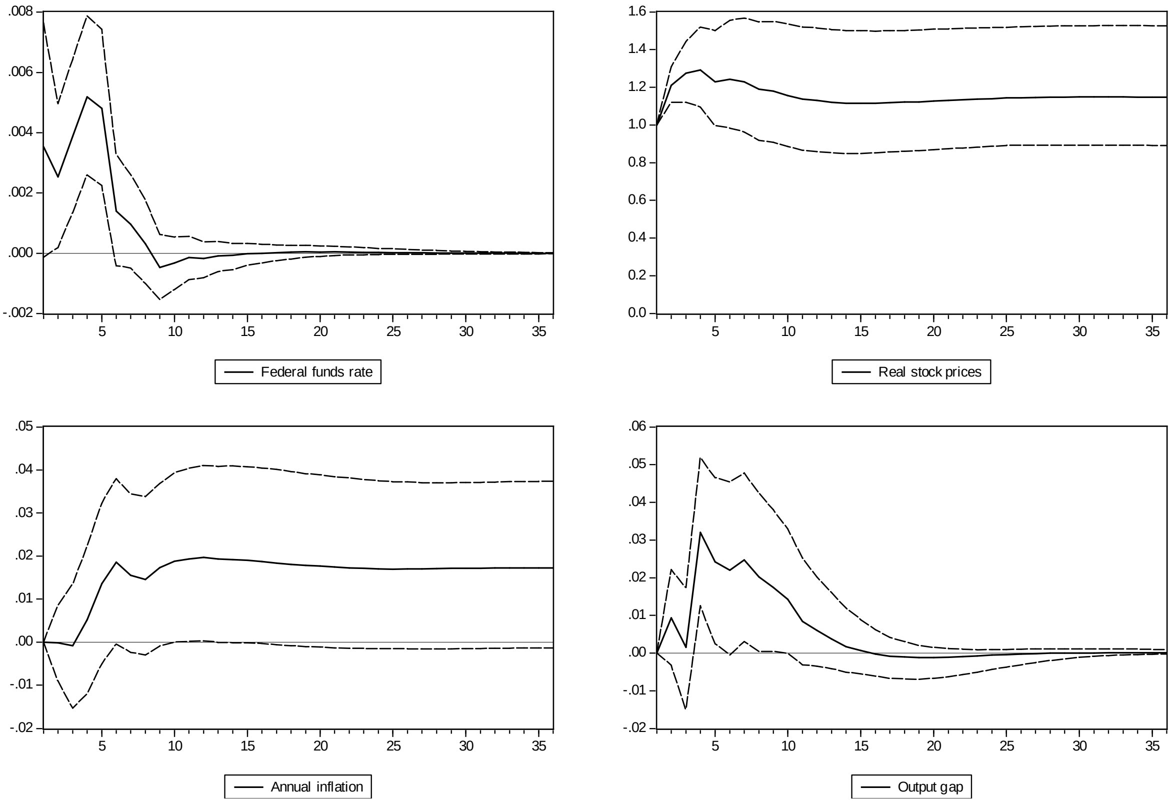
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.21 notes.

Figure A3.27: Impulse responses to FFR shock - without stock market crash dummy



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^a_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The 1-month federal funds futures contract rate (as of last business day of previous month) is included as an exogenous variable. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

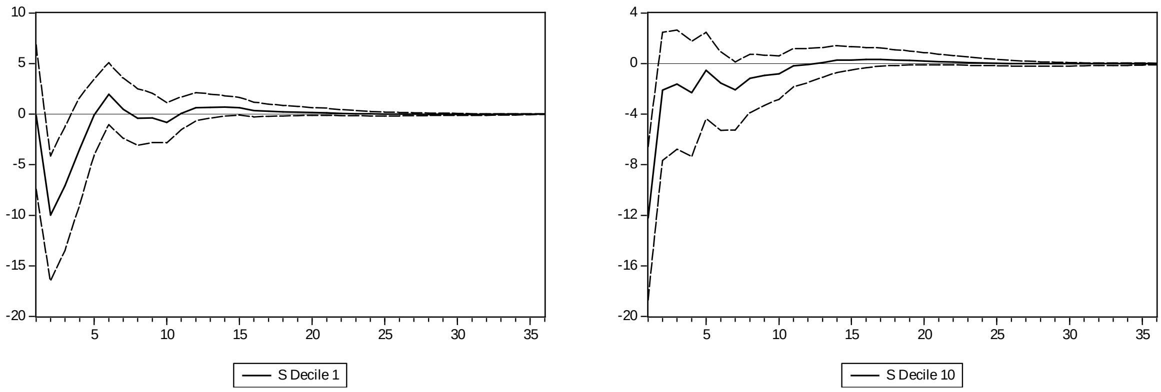
Figure A3.28: Impulse responses to SP shock - without stock market crash dummy



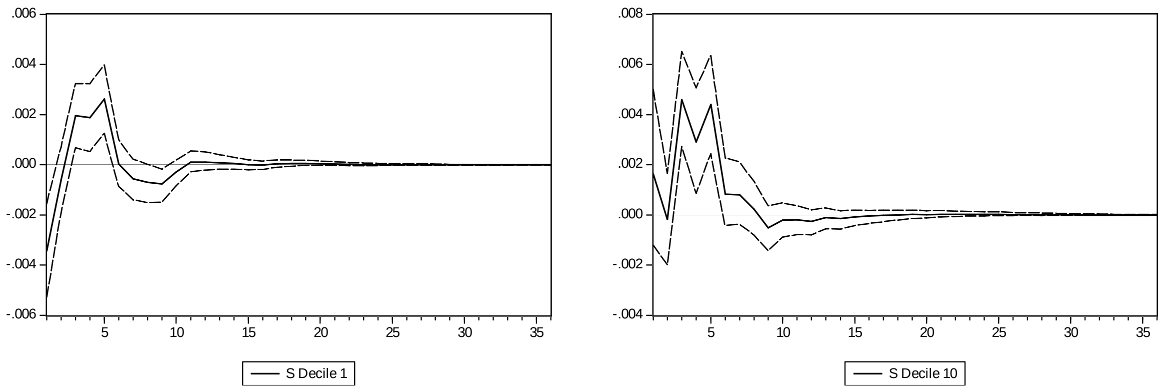
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.27 notes.

Figure A3.29: Impulse responses to FFR and SP shocks - without stock market crash dummy (size decile portfolios)

Panel A: FFR shock



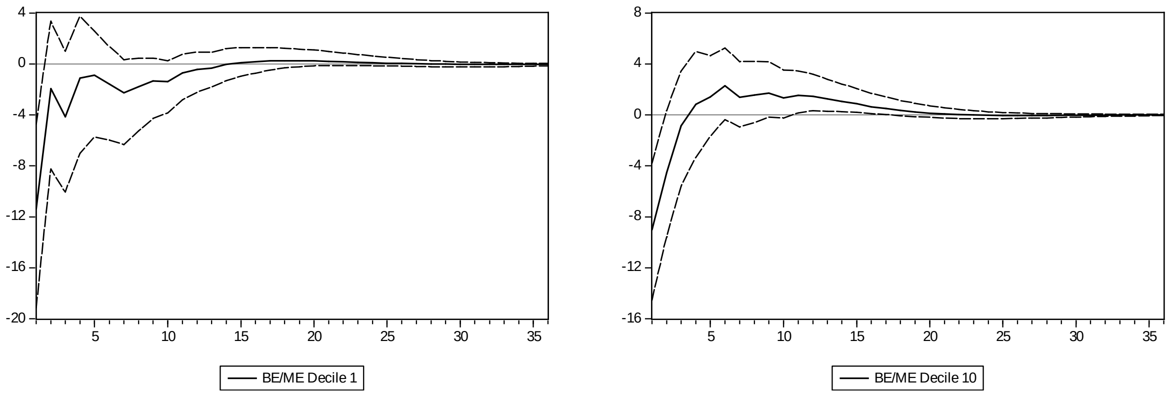
Panel B: SP shock



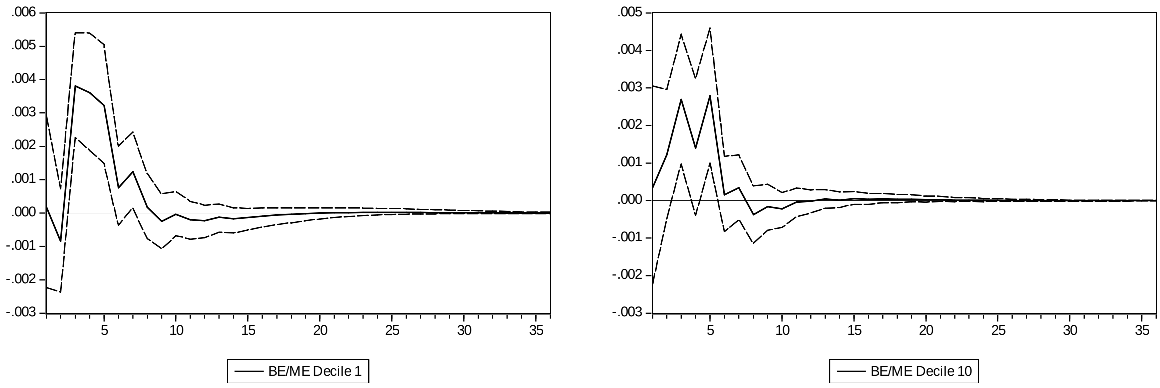
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.27 notes.

Figure A3.30: Impulse responses to FFR and SP shocks - without stock market crash dummy (value decile portfolios)

Panel A: FFR shock

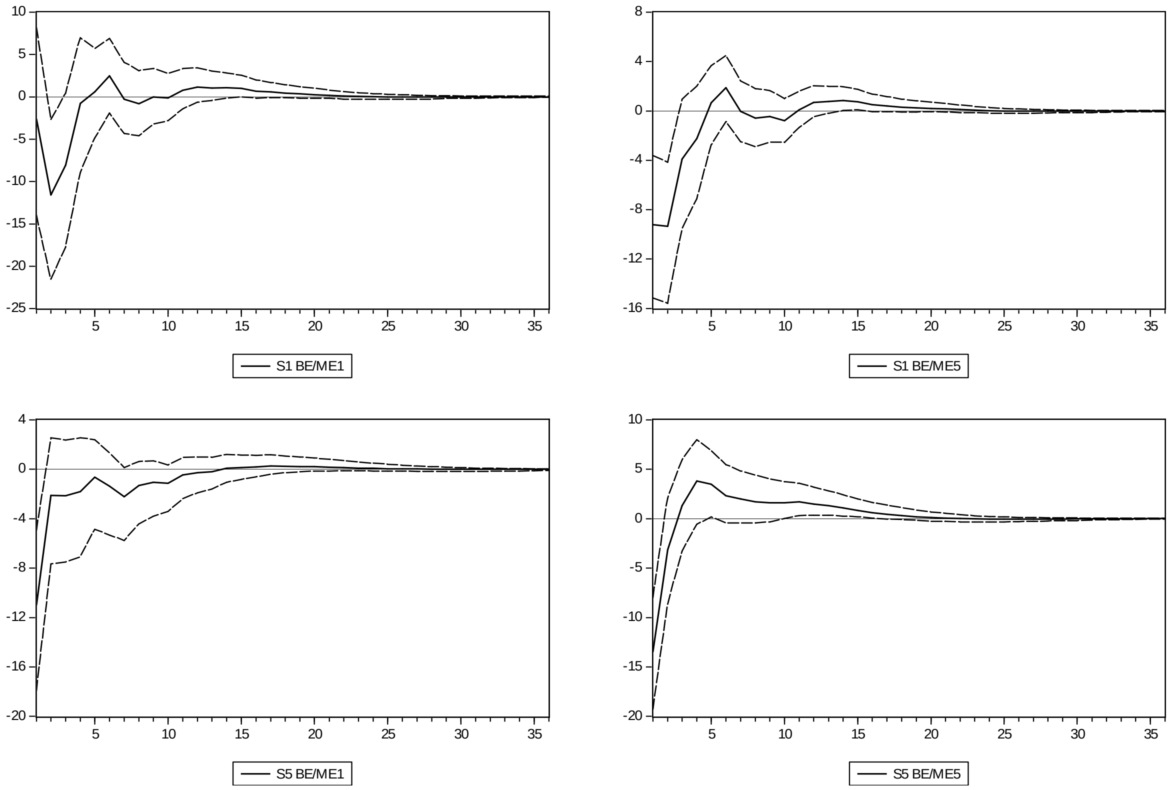


Panel B: SP shock



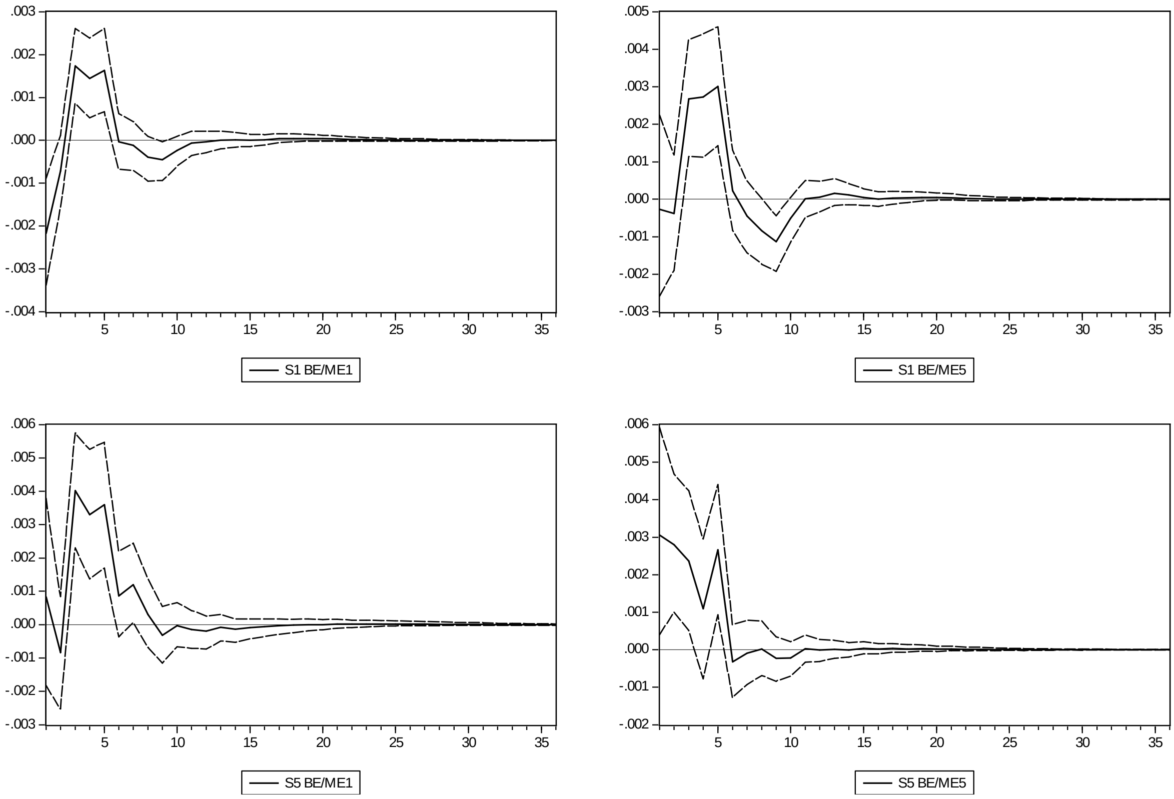
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.27 notes.

Figure A3.31: Impulse responses of real stock returns to FFR shock - without stock market crash dummy (size-value quintile portfolios)



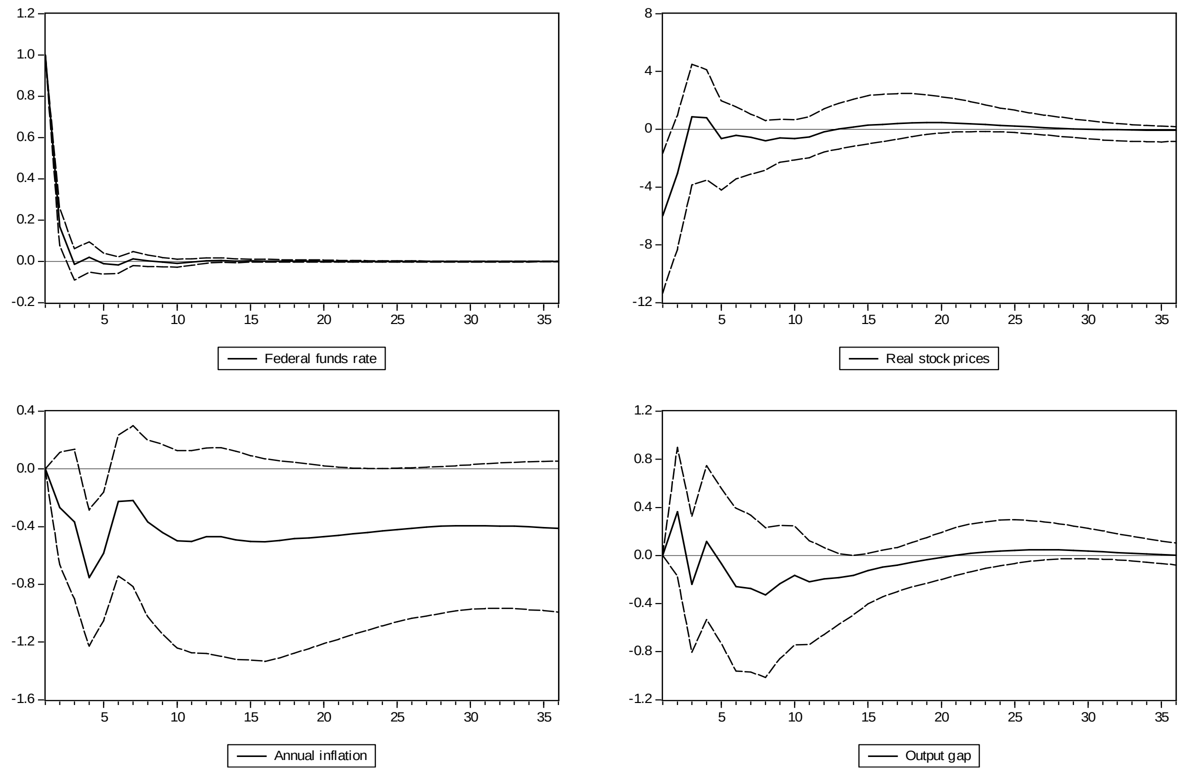
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.27 notes.

Figure A3.32: Impulse responses of the federal funds rate to SP shock - without stock market crash dummy (size-value quintile portfolios)



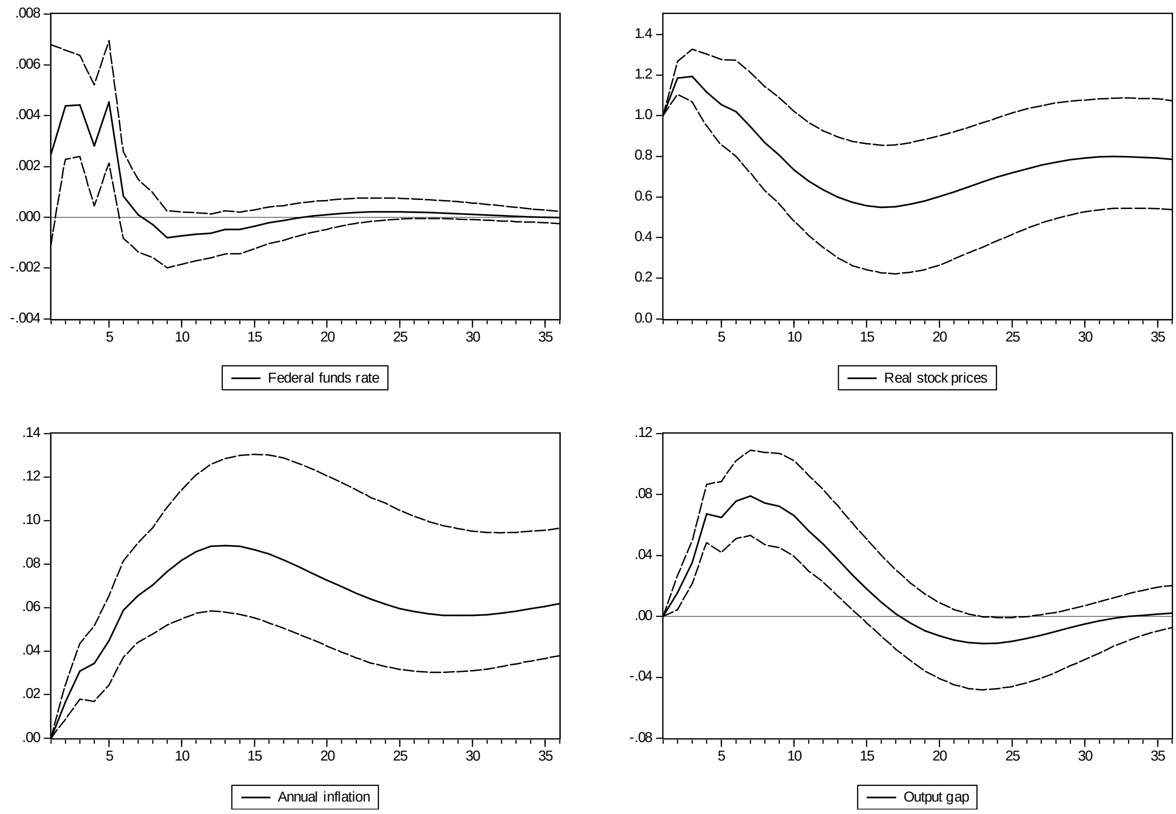
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.27 notes.

Figure A3.33: Impulse responses to FFR shock - extended sample period



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2008:12 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^{a}_{t}$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). Dummy variables that take value of one in 2001:9, 2008:9 and 2008:10, respectively, and zero otherwise, and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

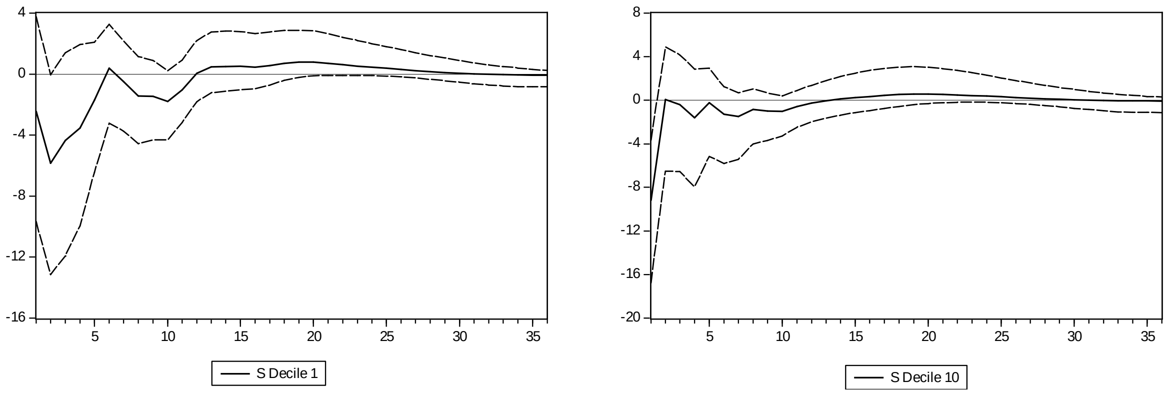
Figure A3.34: Impulse responses to SP shock - extended sample period



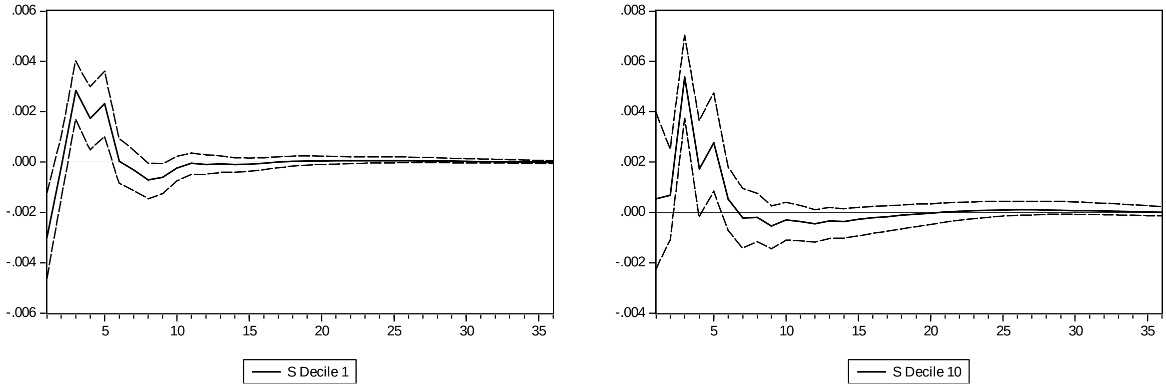
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.33 notes.

Figure A3.35: Impulse responses to FFR and SP shocks - extended sample period (size decile portfolios)

Panel A: FFR shock



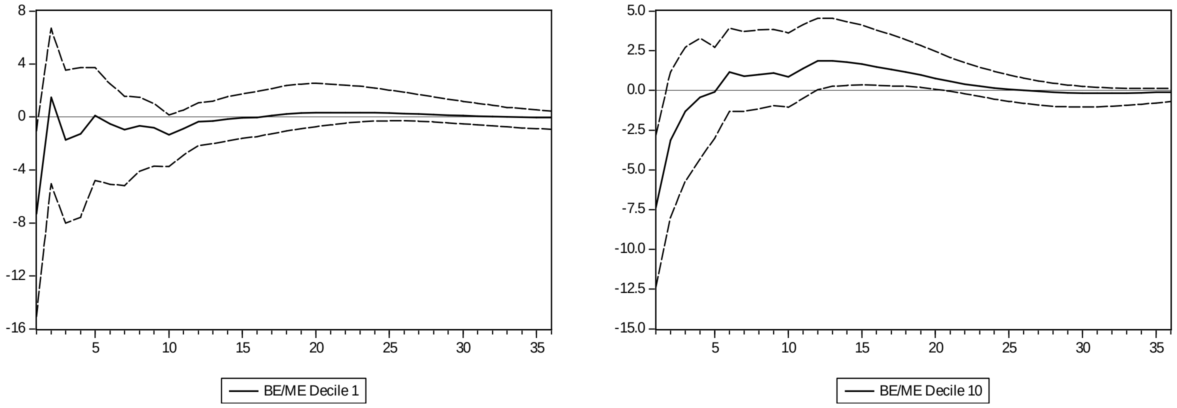
Panel B: SP shock



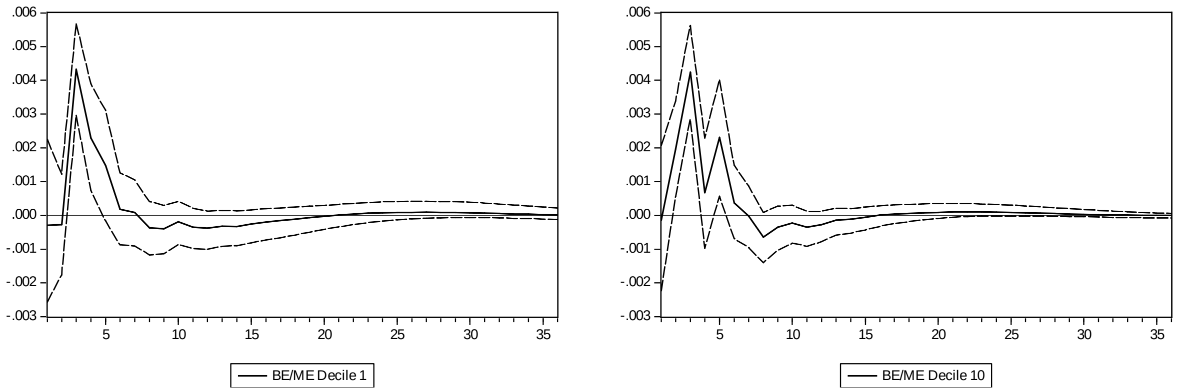
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.33 notes.

Figure A3.36: Impulse responses to FFR and SP shocks - extended sample period (value decile portfolios)

Panel A: FFR shock

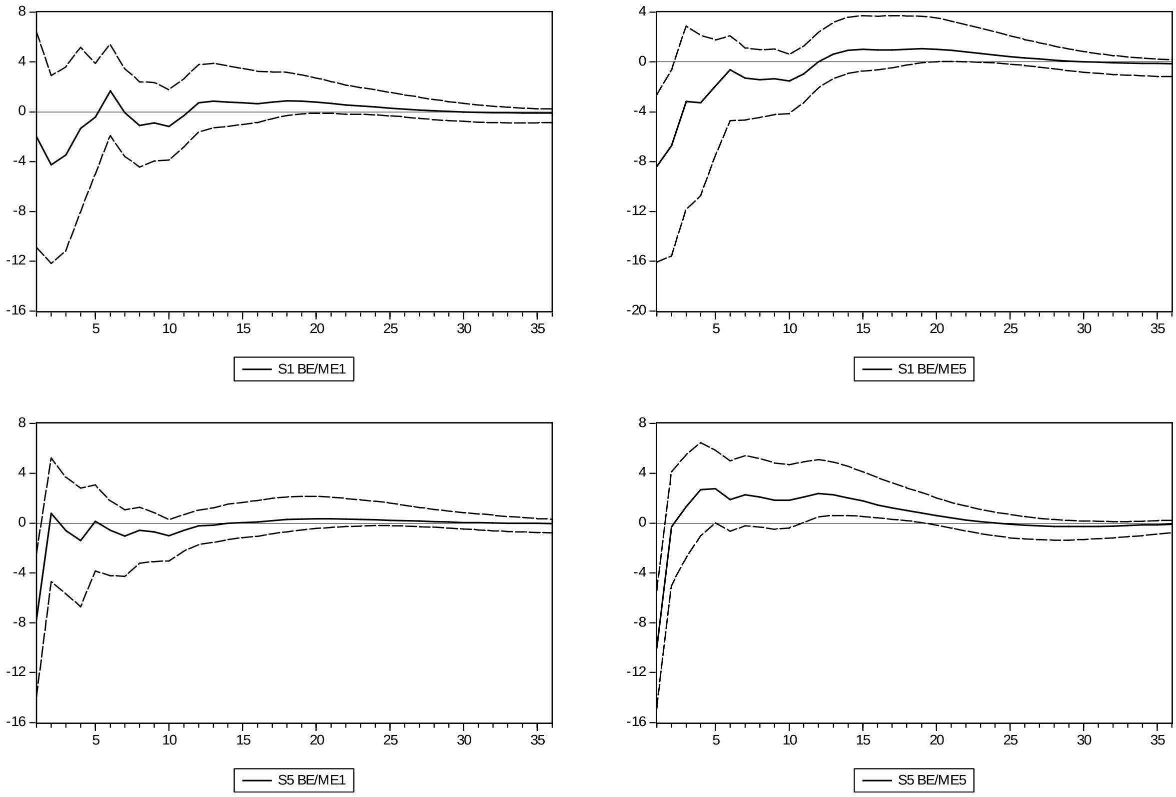


Panel B: SP shock



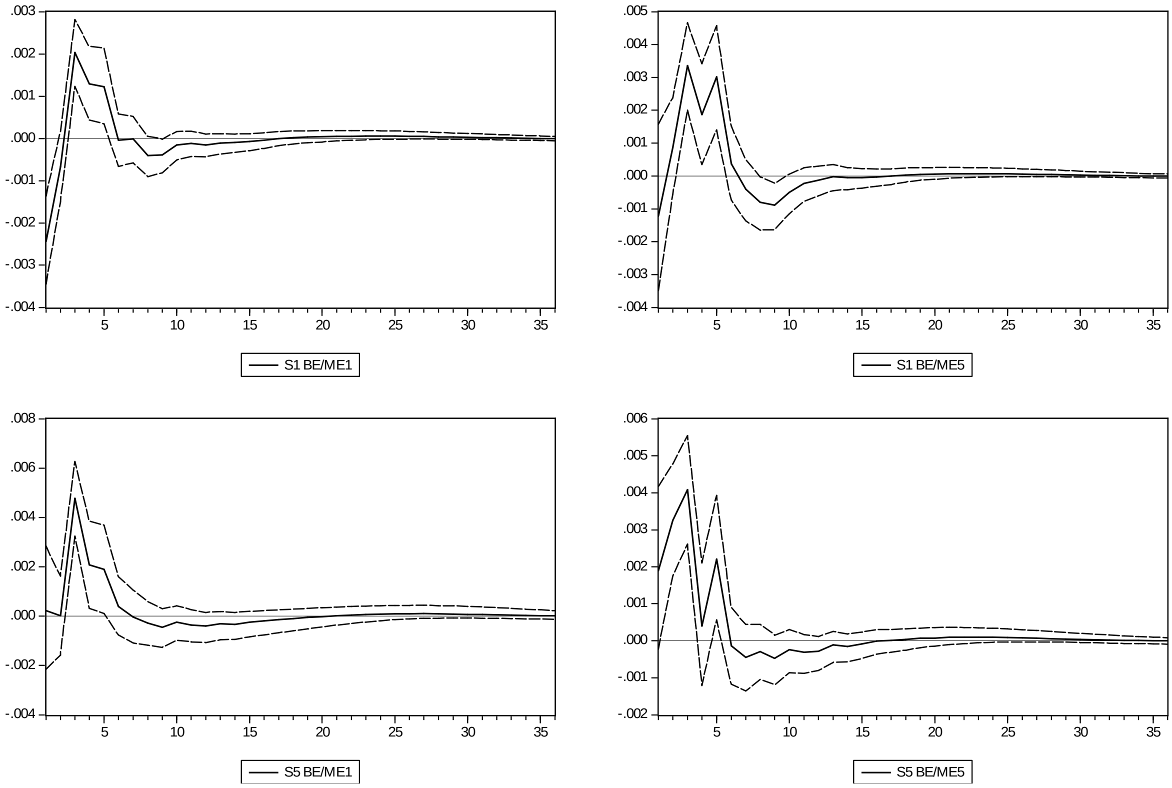
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.33 notes.

Figure A3.37: Impulse responses of real stock prices to FFR shock - extended sample period (size-value quintile portfolios)



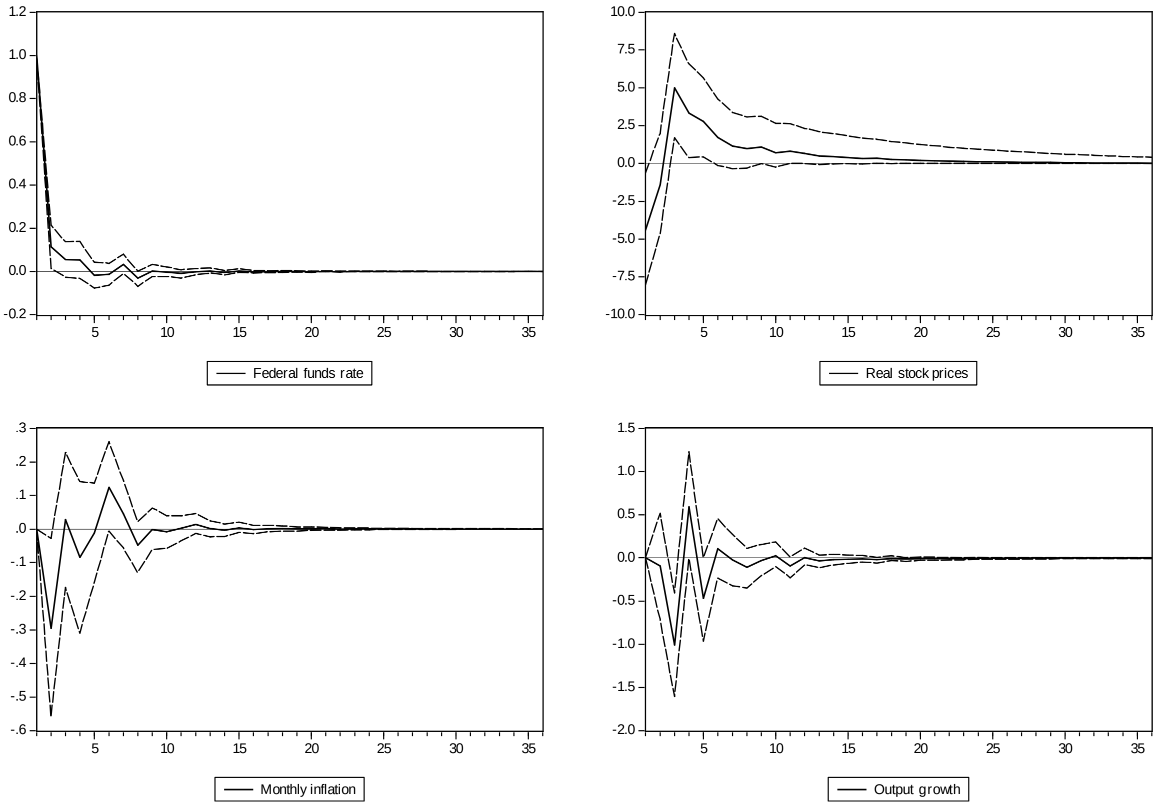
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.33 notes.

Figure A3.38: Impulse responses of the federal funds rate to SP shock - extended sample period (size-value quintile portfolios)



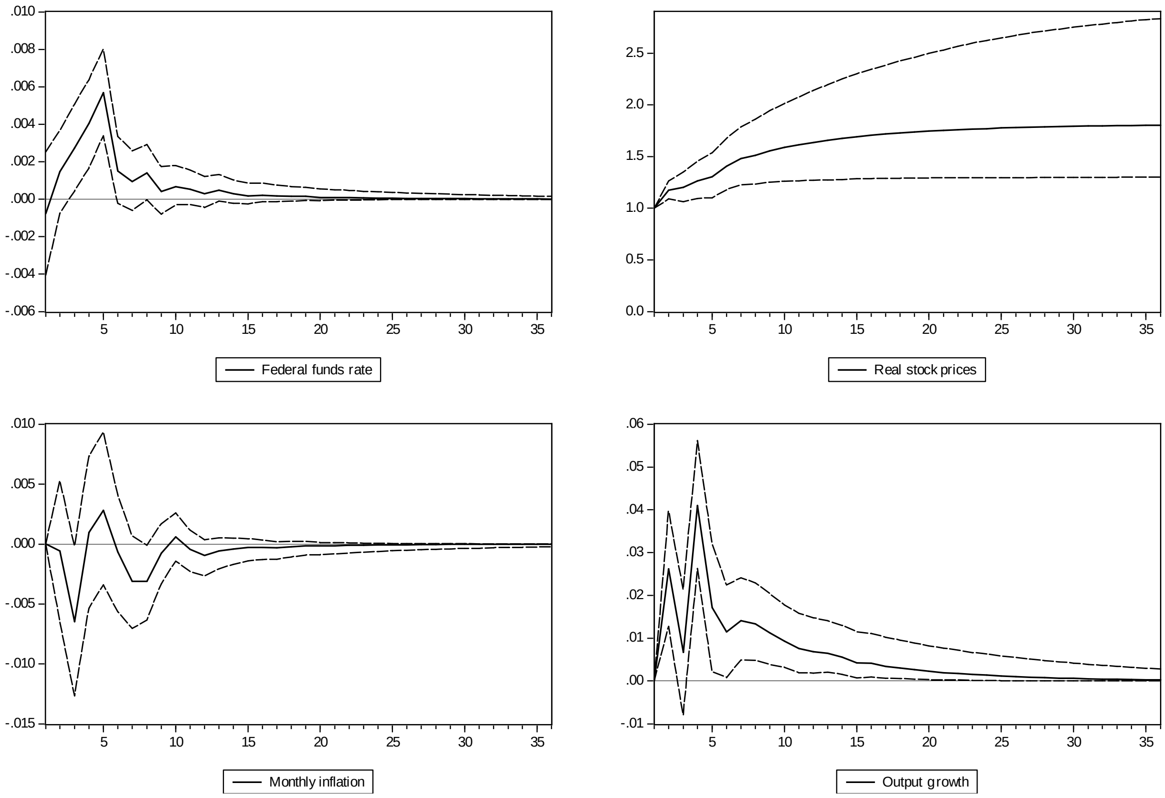
Notes: This figure plots central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-BE/ME quintile portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with relevant real portfolio stock returns. See also Figure A3.33 notes.

Figure A3.39: Impulse responses to FFR shock - alternative data transformation



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, monthly inflation, and monthly industrial output growth following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the monthly change (lagged) in the leading economic indicator (Δlei_{t-1}), the monthly output growth (Δip_t), the monthly inflation (π^m_t), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

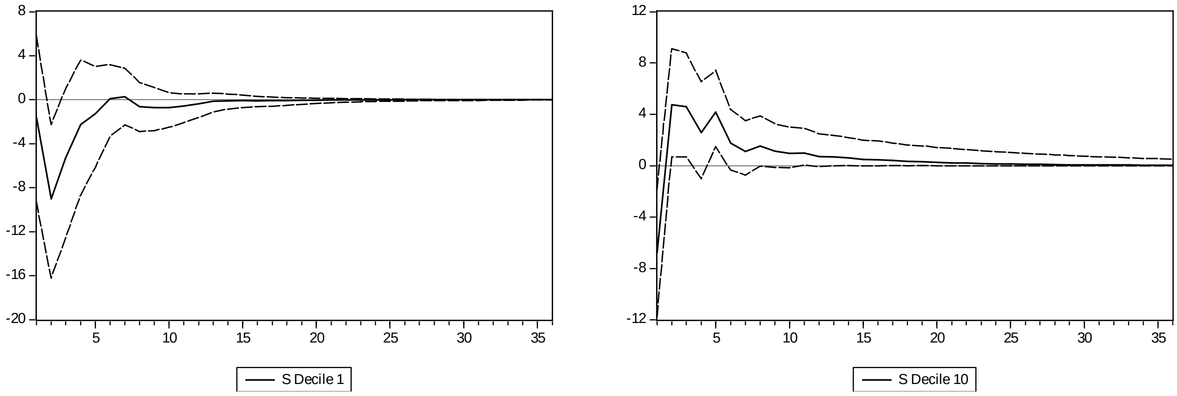
Figure A3.40: Impulse responses to SP shock - alternative data transformation



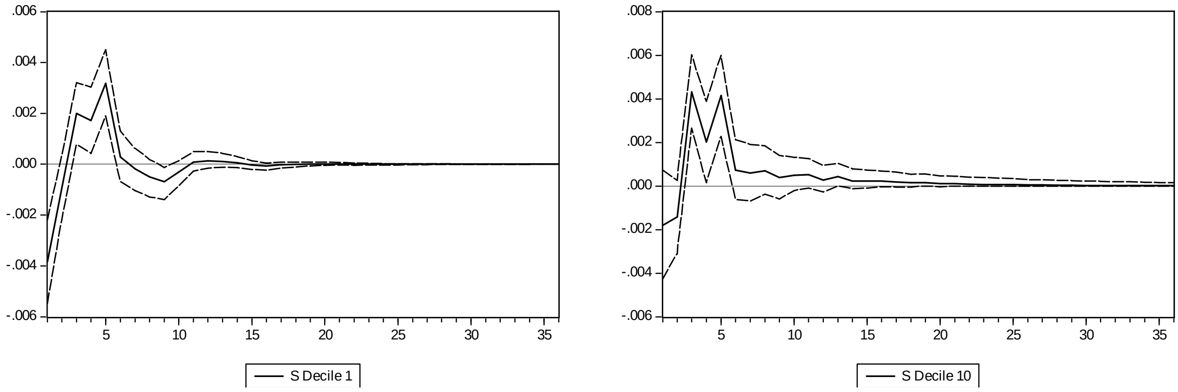
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, monthly inflation, and monthly output growth following a positive 1% stock price shock. See also Figure A3.39 notes.

Figure A3.41: Impulse responses to FFR and SP shocks – alternative data transformation (size decile portfolios)

Panel A: FFR shock



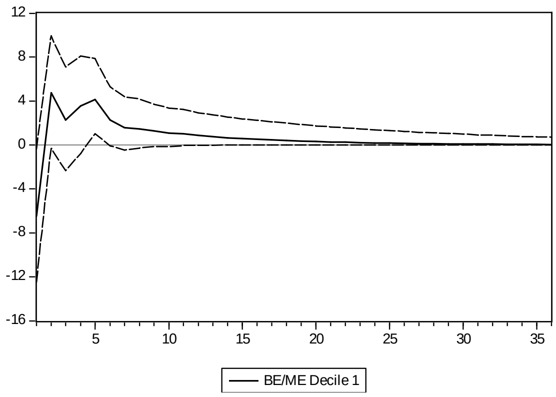
Panel B: SP shock



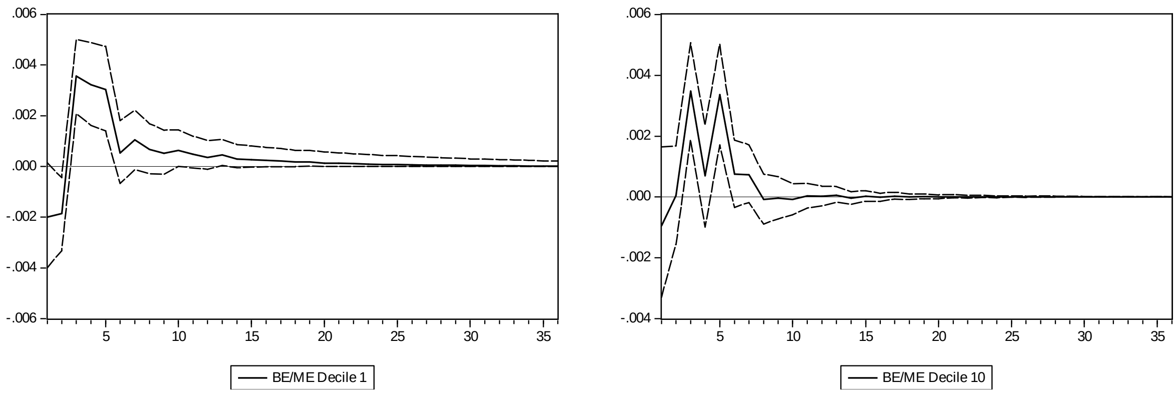
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.39 notes.

Figure A3.42: Impulse responses to FFR and SP shocks - alternative data transformation (value decile portfolios)

Panel A: FFR shock

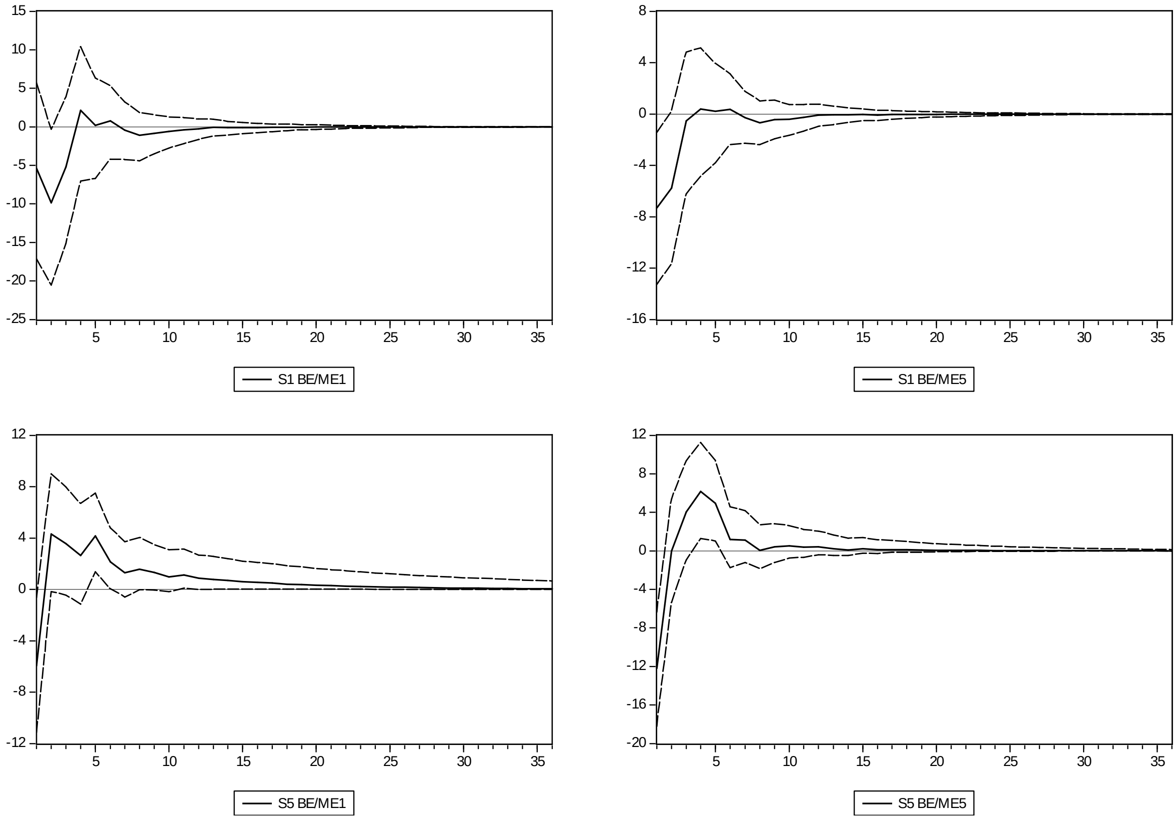


Panel B: SP shock



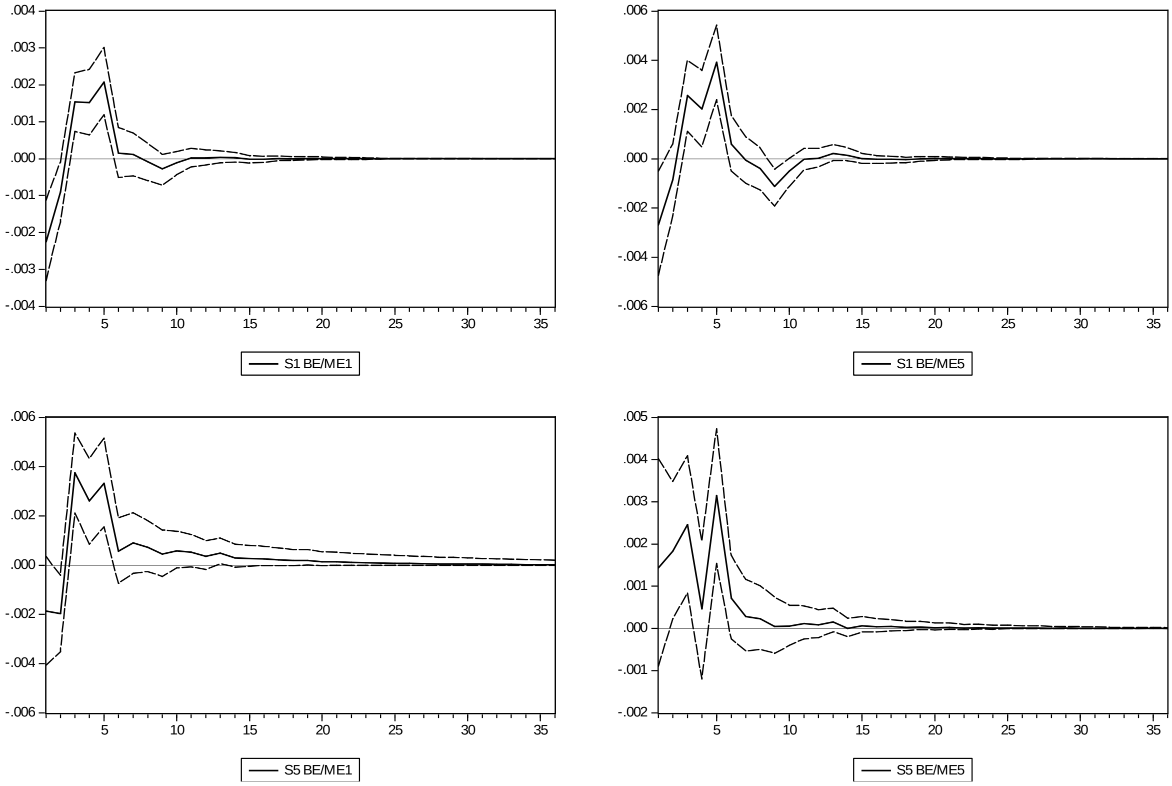
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.39 notes.

Figure A3.43: Impulse responses of real stock prices to FFR shock - alternative data transformation (size-value quintile portfolios)



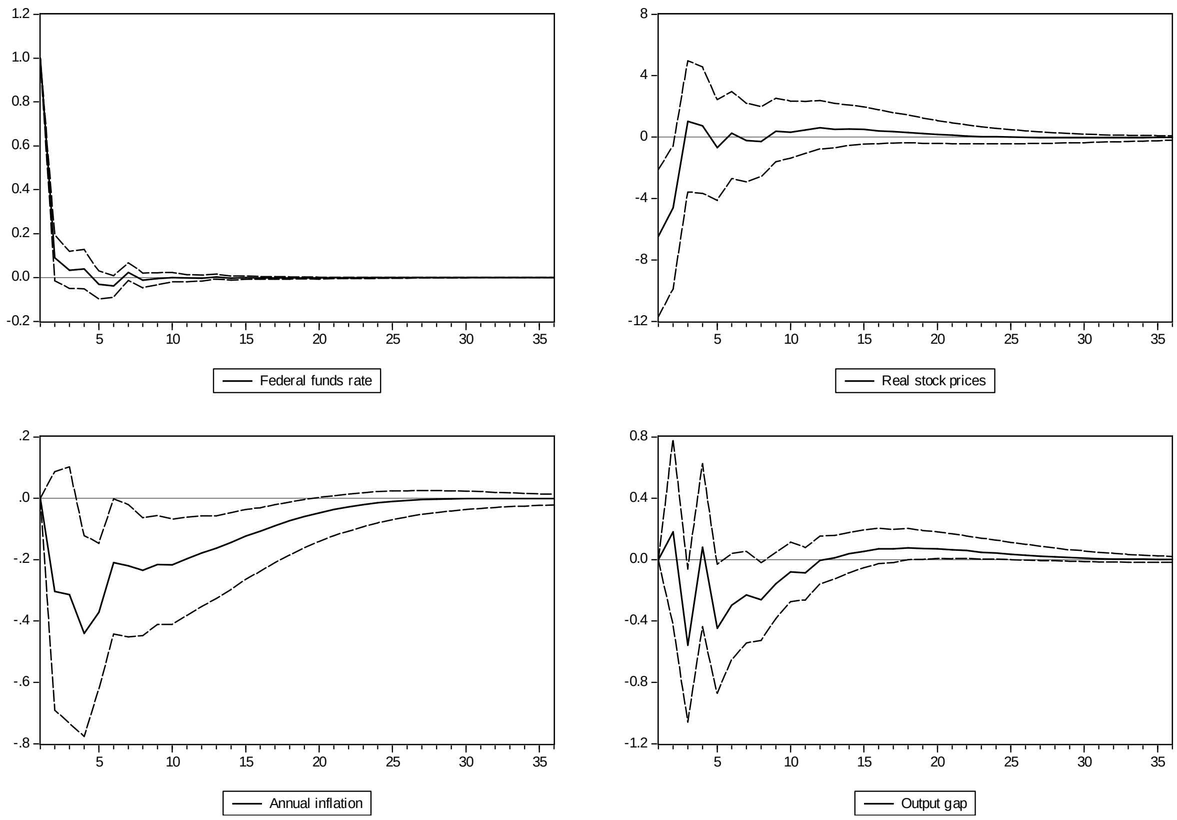
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.39 notes.

Figure A3.44: Impulse responses of the federal funds rate to SP shock - alternative data transformation (size-value quintile portfolios)



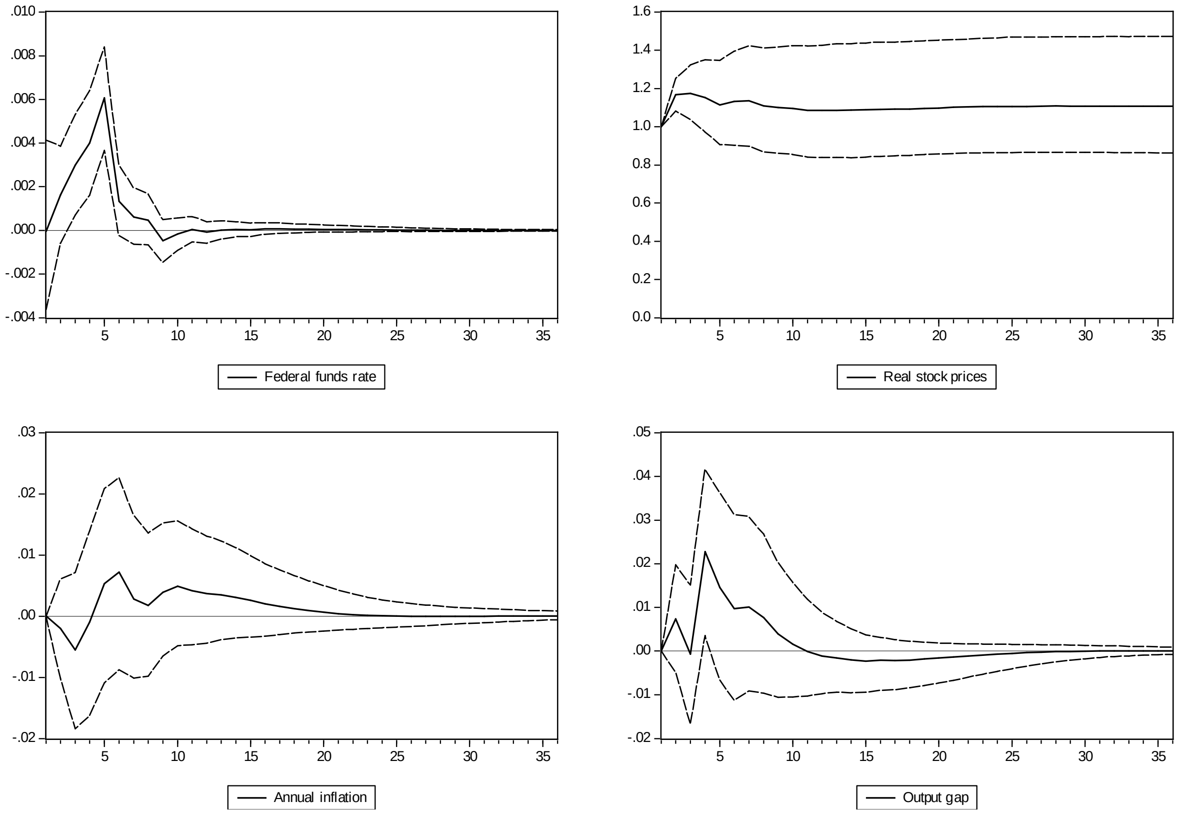
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.39 notes.

Figure A3.45: Impulse responses to FFR shock – undifferenced inflation



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^{a}_{t-1}$), the output gap (gap_t), the annual inflation (π^{a}_{t}), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one in 2001:9 and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

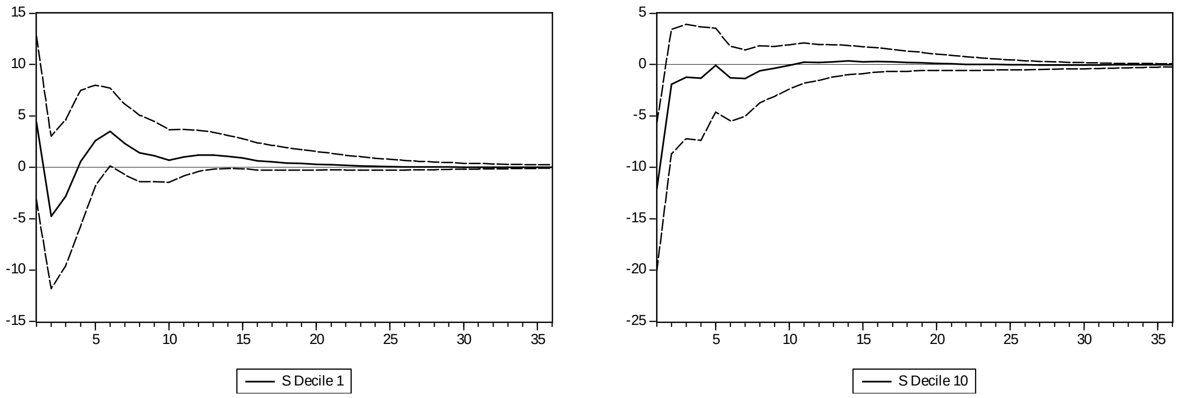
Figure A3.46: Impulse responses to SP shock – undifferenced inflation



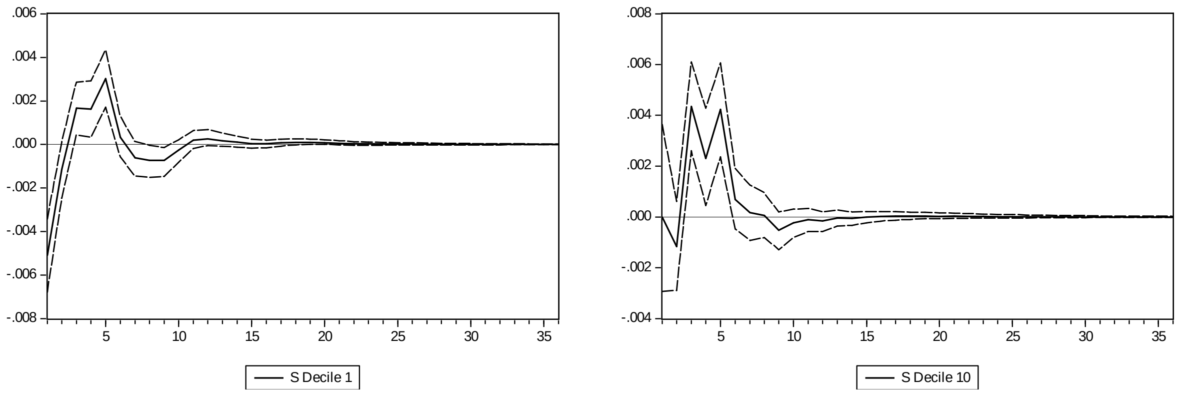
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.45 notes.

Figure A3.47: Impulse responses to FFR and SP shocks – undifferenced inflation (size decile portfolios)

Panel A: FFR shock



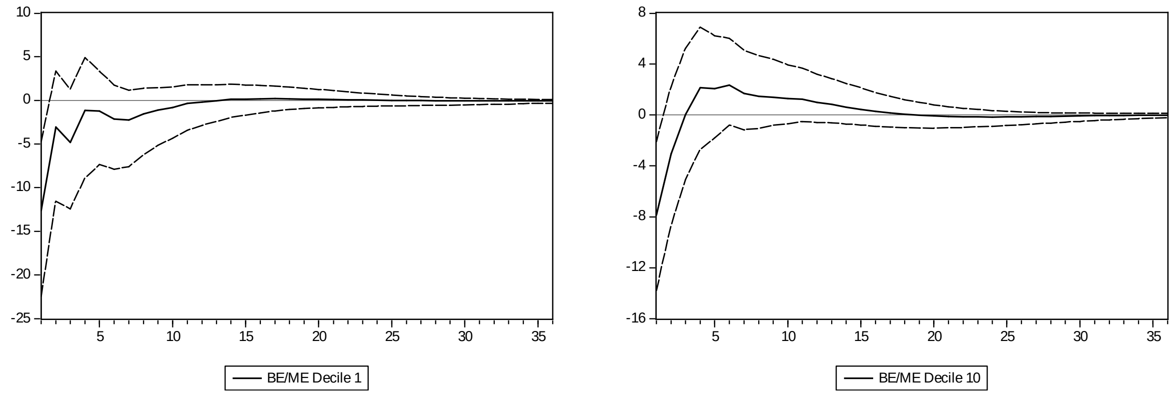
Panel B: SP shock



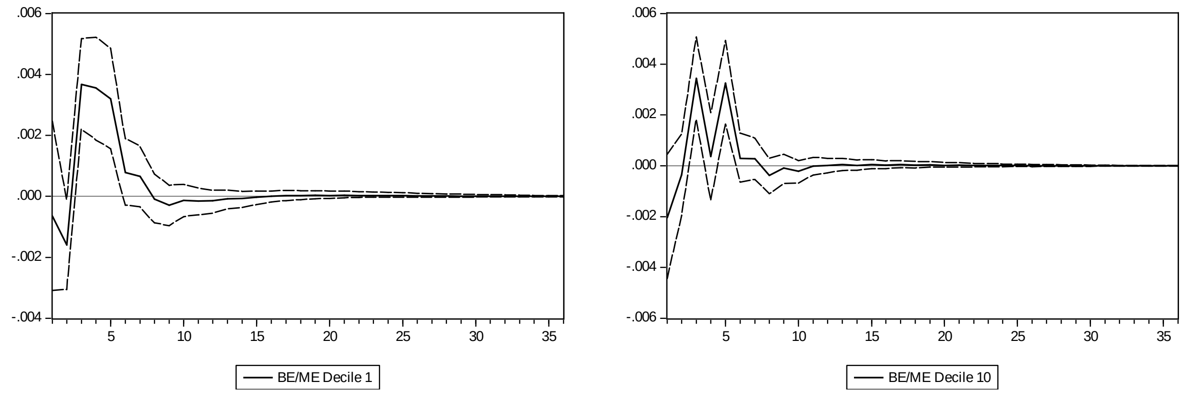
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.45 notes.

Figure A3.48: Impulse responses to FFR and SP shocks – undifferenced inflation (value decile portfolios)

Panel A: FFR shock

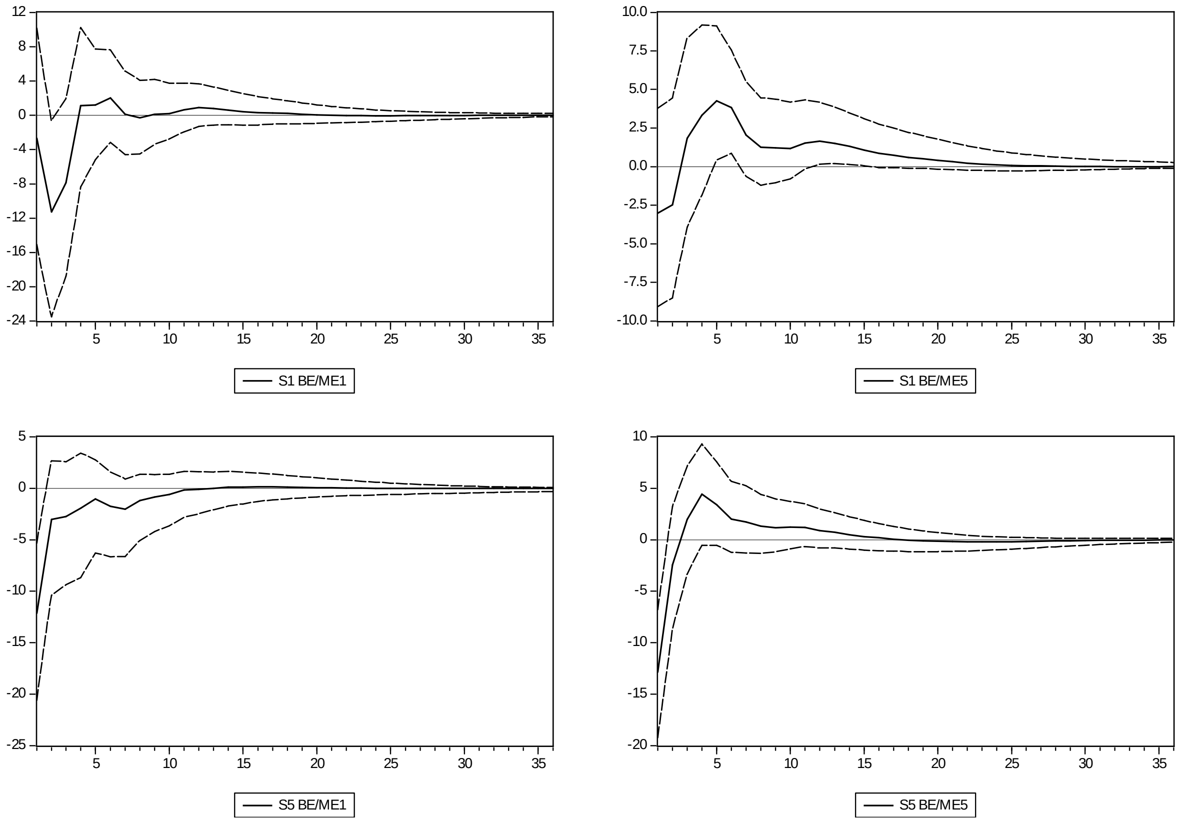


Panel B: SP shock



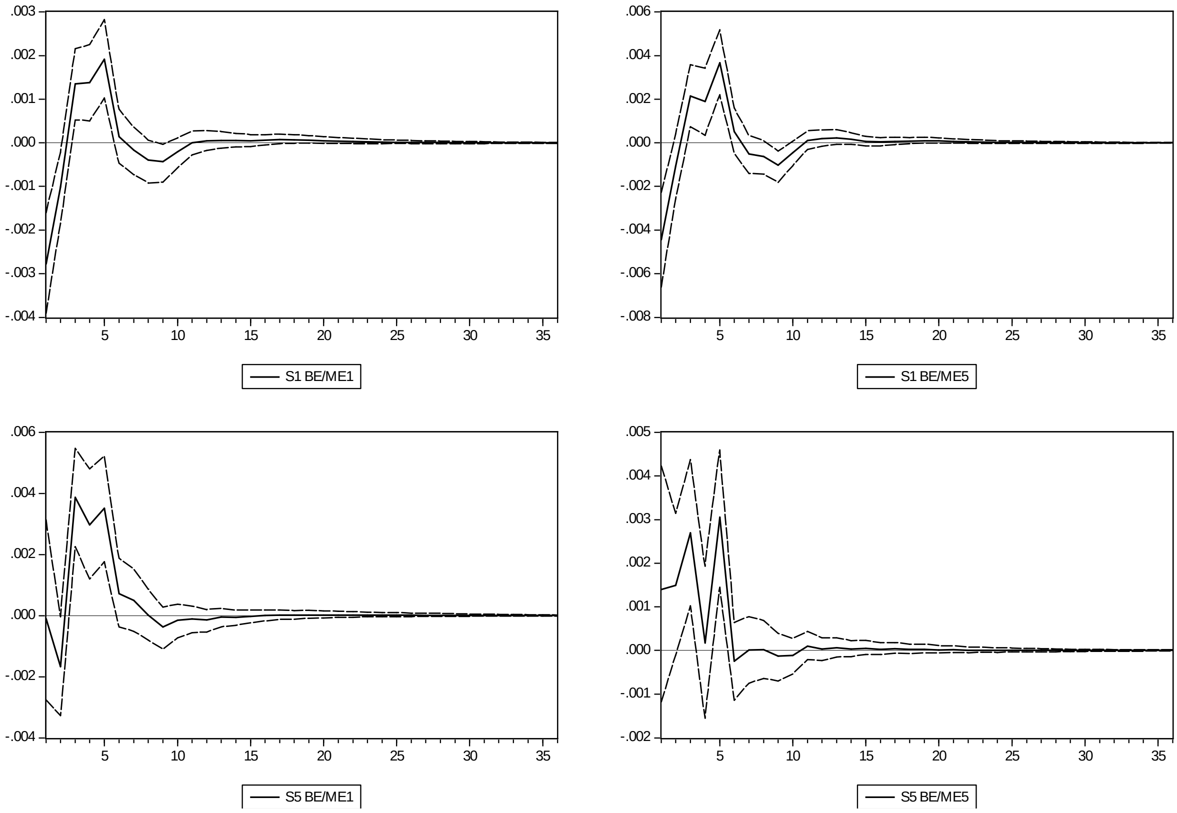
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.45 notes.

Figure A3.49: Impulse responses of real stock prices to FFR shock – undifferenced inflation (size-value quintile portfolios)



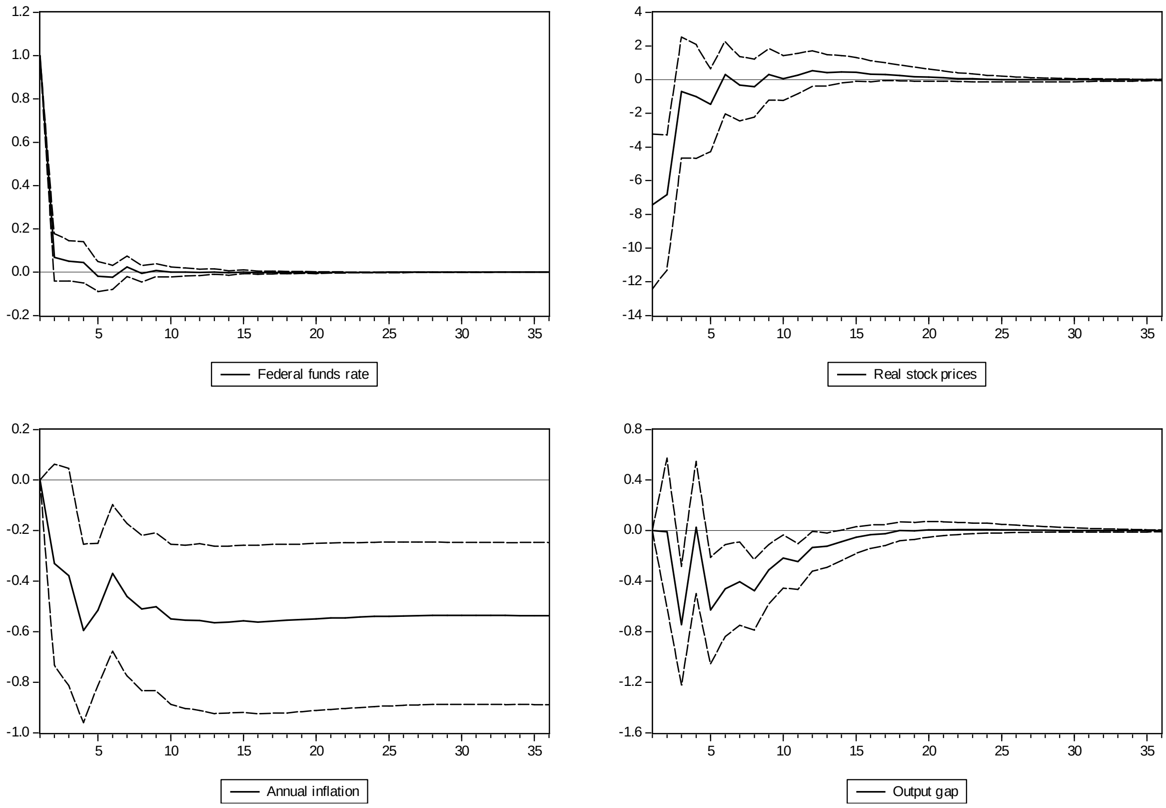
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.45 notes.

Figure A3.50: Impulse responses of the federal funds rate to SP shock – undifferenced inflation (size-value quintile portfolios)



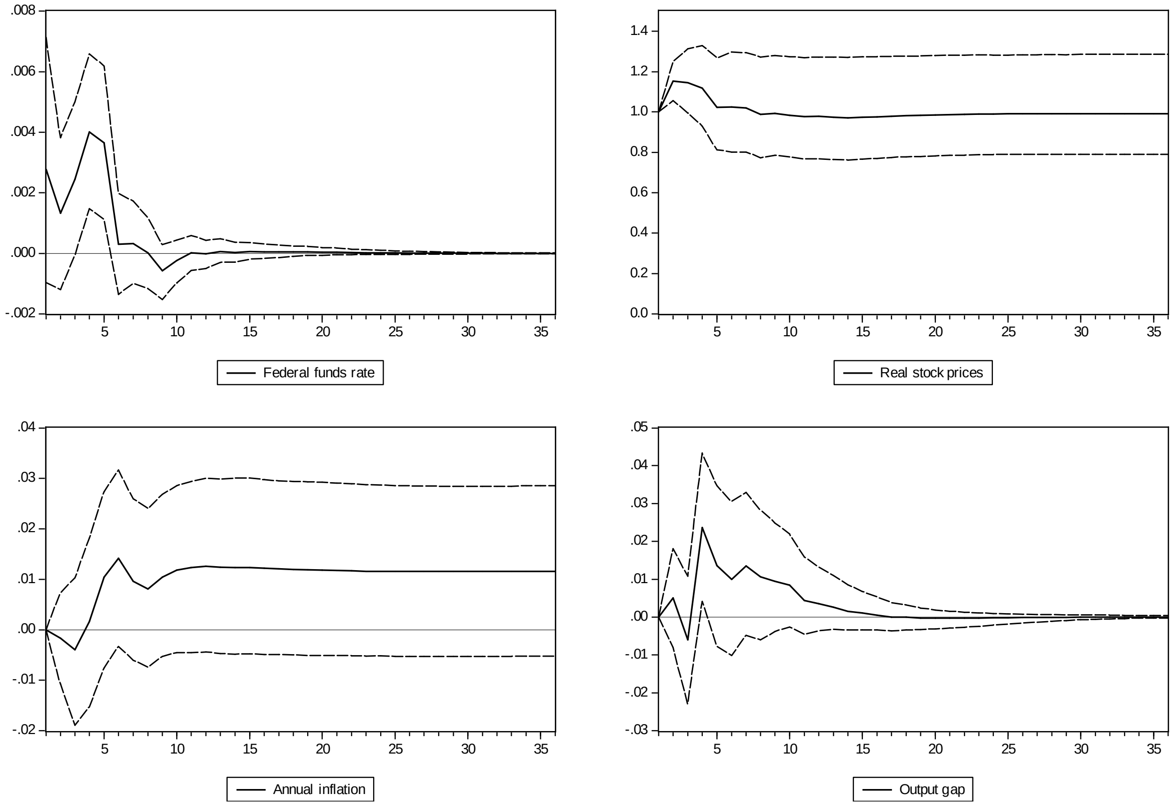
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.45 notes.

Figure A3.51: Impulse responses to FFR shock – US recession dummy



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated over the sample period 1994:2 – 2007:7 including 4 lags. The state vector contains the first difference (lagged) of annual change in the leading economic indicator ($\Delta lead^a_{t-1}$), the output gap (gap_t), the first difference of annual inflation ($\Delta\pi^a_t$), the monthly real stock market returns (Δsp_t) and the federal funds rate (i_t). The dummy variable that takes value of one during US recessions and zero otherwise and the 1-month federal funds futures contract rate (as of last business day of previous month) are included as exogenous variables. The dashed lines represent 68% Bayesian probability bands generated using Monte Carlo integration with 10000 draws as suggested by Doan (2015).

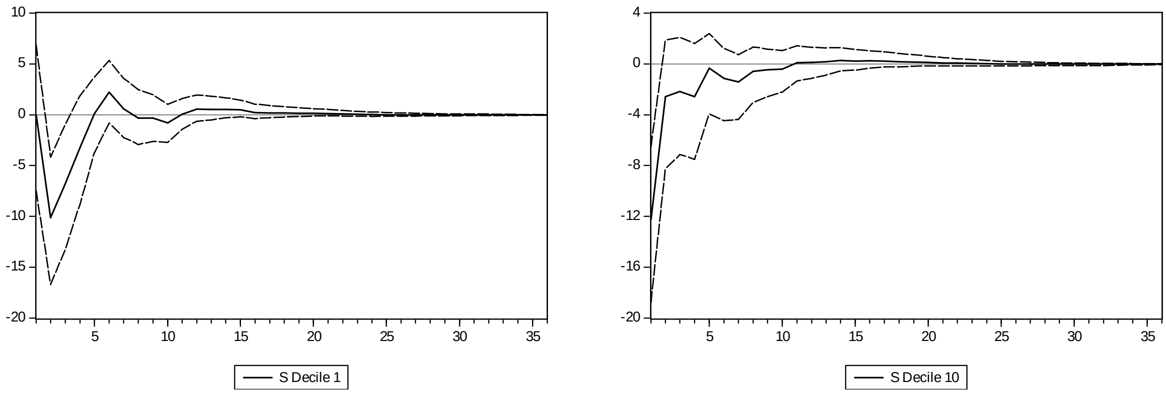
Figure A3.52: Impulse responses to SP shock – US recession dummy



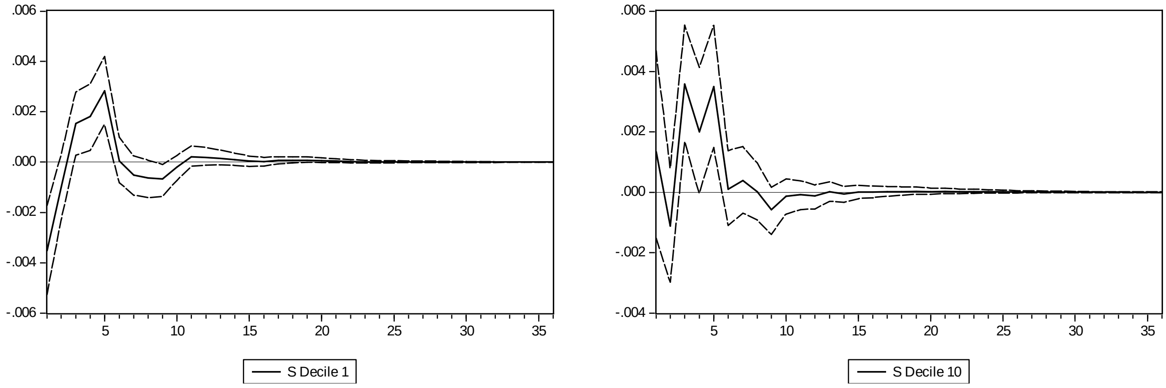
Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate, real stock prices, annual inflation, and output gap following a positive 1% stock price shock. See also Figure A3.51 notes.

Figure A3.53: Impulse responses to FFR and SP shocks – US recession dummy (size decile portfolios)

Panel A: FFR shock



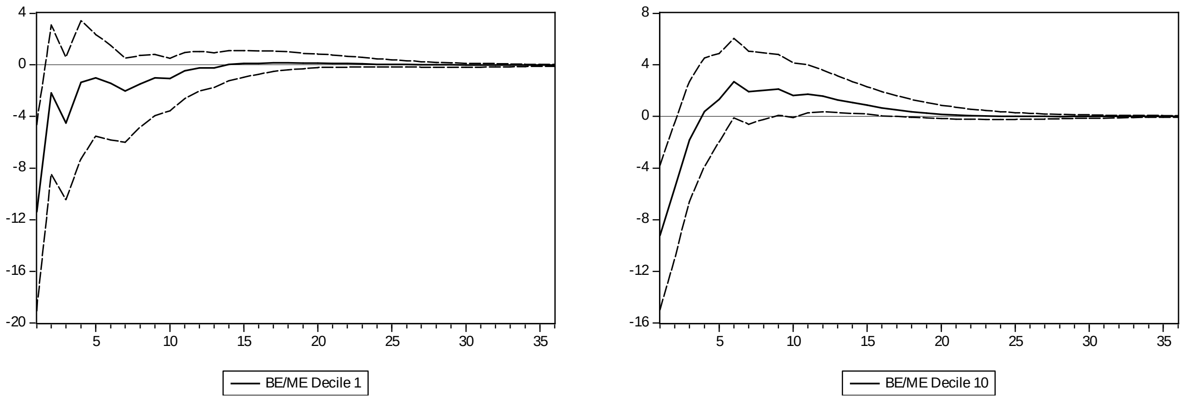
Panel B: SP shock



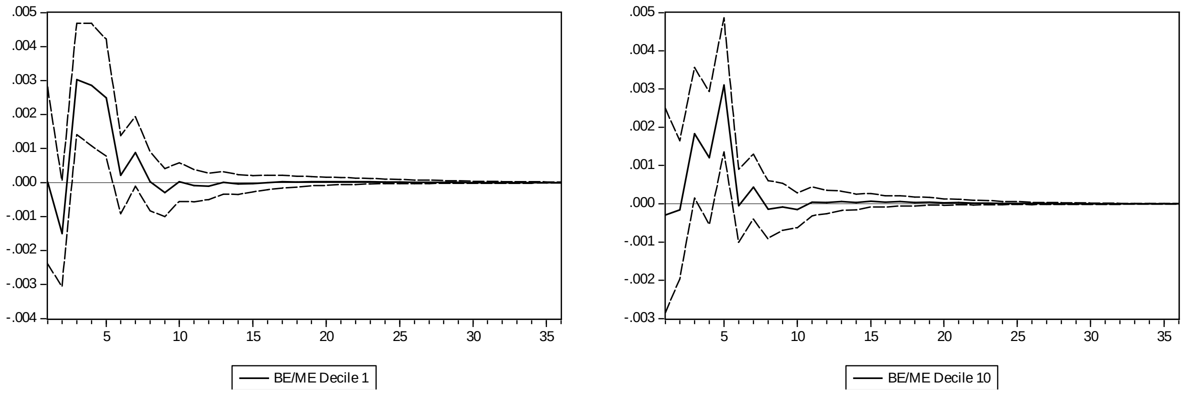
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile size-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile size-sorted portfolios (Panel B). The augmented SVAR model is estimated for each size decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.51 notes.

Figure A3.54: Impulse responses to FFR and SP shocks – US recession dummy (value decile portfolios)

Panel A: FFR shock

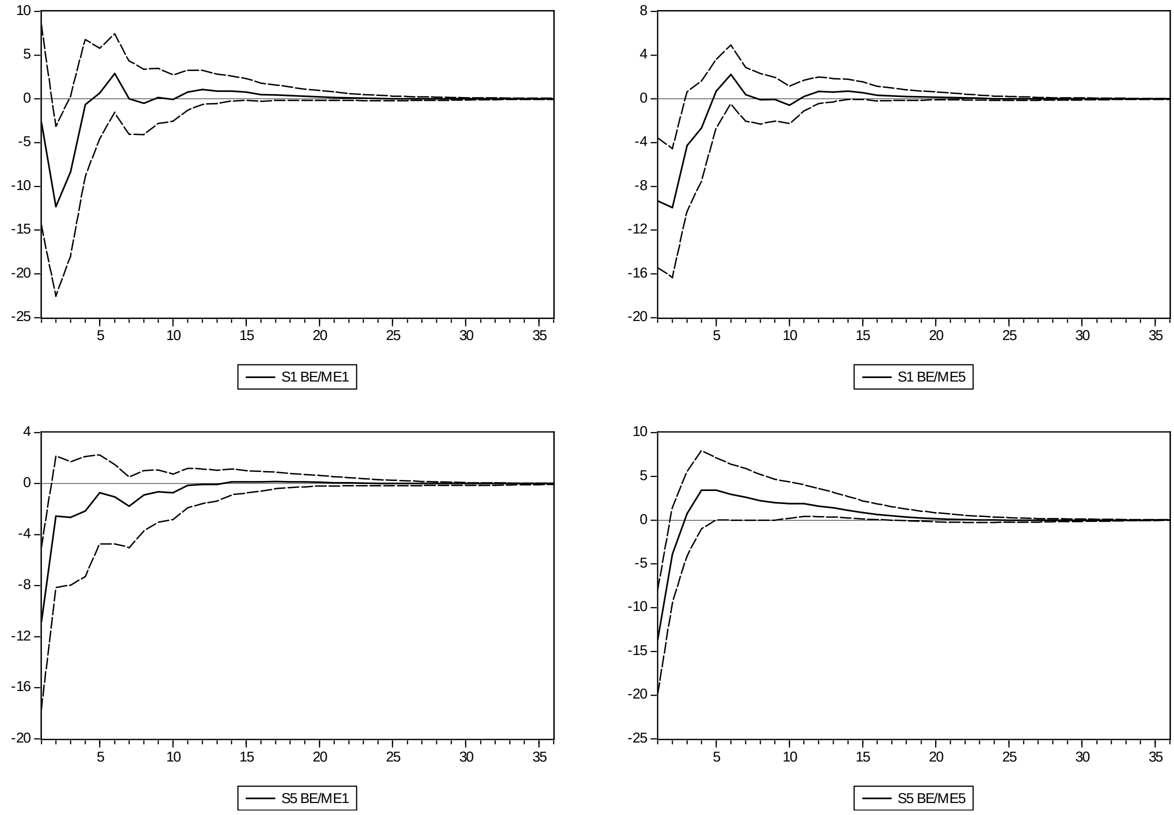


Panel B: SP shock



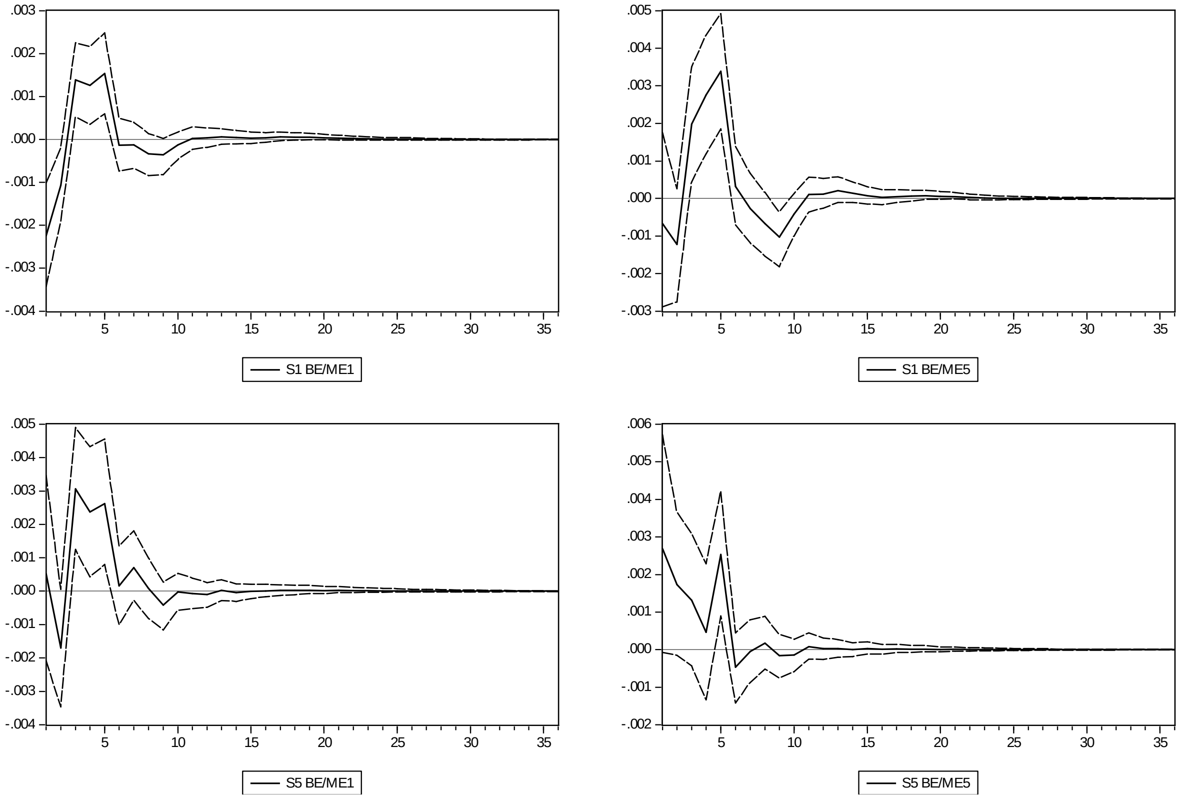
Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices in the first and tenth decile value-sorted portfolios following a 1-percentage-point contractionary monetary policy shock (Panel A) and for the federal funds rate following a positive 1% stock price shock in the first and tenth decile value-sorted portfolios (Panel B). The augmented SVAR model is estimated for each value decile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.51 notes.

Figure A3.55: Impulse responses of real stock prices to FFR shock - US recession dummy (size-value quintile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for real stock prices (sp_t) across the selected double-sorted size-value portfolios (intersections between extreme size and value quintiles) following a 1-percentage-point contractionary monetary policy shock. The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.51 notes.

Figure A3.56: Impulse responses of the federal funds rate to SP shock - US recession dummy (size-value quintile portfolios)



Notes: This figure plots the central measures of impulse responses (median, solid line) for the federal funds rate following a positive 1% stock price shock across selected double-sorted size-value portfolios (intersections between extreme size and value quintiles). The augmented SVAR model is estimated for each size-value quintile portfolio by replacing real stock market returns in the state vector with the relevant portfolio real stock returns. See also Figure A3.51 notes.

Conclusion

The financial system is the bloodline of the economy and severe disruptions to the functioning of financial markets may endanger price stability and economic growth. The global financial crisis of 2007 – 2009 not only had financial and economic implications worldwide, but it also left academics and monetary policymakers with plenty of interesting and challenging research questions to answer. Asset prices play an important role in the transmission of monetary policy to aggregate economy and may also provide information about future financial instability. While the literature documents the evidence that monetary policy related news is a significant determinant of asset prices, there is also a rationale for monetary authorities to consider financial market developments when setting the policy rate in addition to standard macroeconomic variables as suggested by Taylor (1993).

The debate of the late 1990s on the appropriate monetary policy response to financial market developments has been rekindled and has led to a surge of theoretical studies analysing whether financial indicators should or should not be considered by monetary authorities. The pre-crisis consensus appears to have shifted towards the view that financial imbalances should be addressed by monetary policymakers over and above their effect on inflation and output forecasts. Furthermore, the empirical evidence on the policy response to financial markets has been revisited by many and special attention has been paid to the global financial crisis. The findings regarding the direct reaction of the Federal Reserve to asset prices and financial conditions are somewhat mixed. On the other hand, it seems to be largely agreed that the financial crisis has changed the way monetary policy is conducted for at least that period. Following the zero lower bound on a short-term nominal interest rate and the adoption of unconventional monetary policy tools, there has been a surge in the literature that examines the impact of central bank’s asset purchases on longer-term interest rates and other asset prices as well as on macroeconomy. In general, the financial crisis has reshaped the pre-crisis thinking and highlighted the importance of financial imbalances for real economy and, thus, policy making.

The motivation for the empirical analyses of this thesis stems from developments in financial markets and the response of monetary policy to these developments over the period 2007 - 2014. This doctoral thesis revisits the relationship between monetary policy conduct and financial market developments in the US over the period of Great Moderation, the global financial crisis and its aftermath. Combined together, three empirical chapters provide the in-depth study of the role of asset prices and financial instability in the monetary policy reaction function and the impact of conventional and unconventional monetary policy actions on the pricing of government bonds and stocks.

Motivated by events around 2007 – 2008, Chapter 1 examines the impact of financial market stress on setting the monetary policy interest rate. The results indicate the direct Fed's reaction to developments in the stock market index, the interest rate (credit) spread, the measure of stock market liquidity and broad financial conditions captured by the financial conditions index. Nevertheless, it is also demonstrated that this reaction is strongly dependent on the business cycle. Specifically, financial market developments have much more weight on the Fed's interest rate decisions in economic recessions as compared to the periods of economic expansions. Moreover, it appears that this significant reaction during economic recessions can be explained, to a large extent, by the Fed's actions in response to the global financial crisis. With respect to the indirect reaction, the Fed's response to inflation declines to some extent and the output gap parameter becomes statistically insignificant in light of elevated financial distress. Nevertheless, the indirect response to financial market stress strengthens in 2007 - 2008. The parameter on expected inflation declines significantly, turns negative and statistically insignificant. With respect to the output gap, the estimated coefficient increases slightly, but not substantially, and remains significant. Overall, the finding that worsening financial conditions imply a lower policy rate is largely driven by the Fed's actions in the period 2007 - 2008. Hence, the latest crisis had a significant impact on the Fed's monetary policy framework at that time with the focus shifting away from price stability and, possibly, more towards the smooth functioning of the financial system and financial stability.

Chapter 2 investigates empirically what explains the variation in unexpected excess returns on the 2-, 5- and 10-year US Treasury bonds and how the bond market responds to conventional and unconventional monetary policy shocks. The main findings show that the revisions in rational expectations about future inflation is the key driver of the total variability of unexpected excess Treasury bond returns across different maturities. In general, monetary easing is associated with higher unexpected excess Treasury bond returns, i.e. lower bond yields. Furthermore, the results highlight the prominent role of the inflation news component in explaining conventional and unconventional monetary policy effects on bond returns. The positive effect of monetary policy easing on unexpected excess Treasury bond returns is largely explained by a corresponding negative effect on inflation expectations. This implies that the evidence provided in Chapter 2 does not support the portfolio balance channel of quantitative easing as there is no strong response of risk premium news to monetary base expansion. Nevertheless, it is also found that the bond returns reaction to conventional policy shocks has become weaker since the middle 1990s, possibly reflecting changes in the implementation and communication of the Fed's policy. Meanwhile, the results with respect to the quantity-based monetary policy indicators, i.e. unconventional monetary policies, are driven to a great extent by the peak of the financial crisis in autumn of 2008.

As the Federal Reserve has started to normalise monetary policy, Chapter 3 goes back to examining conventional monetary policy. The main empirical analysis is focused over the period of relatively stable economic and financial conditions accounting for significant changes in the Fed's communication to the public. It investigates the impact of interest rate-based monetary policy shocks on real stock prices at aggregate market and portfolio levels taking fully into account the potential simultaneous interaction between the policy rate and real stock returns.

The results confirm a strong, negative and significant monetary policy tightening effect on real stock prices. The findings based on stock portfolios provide the insight to the differential response of stocks to monetary policy as implied by the credit channel of monetary transmission. The empirical evidence highlights the delayed size effect of monetary policy shocks. Following an unexpected increase in the federal funds rate, the initial decline in stock prices of large firms is more pronounced as compared to small firms. In the second period after the shock, however, large stocks recover to a great extent, while small stocks drop sharply. This may be explained through the relative illiquidity and less frequent trading of smaller stocks or through the liquidity pull-back and portfolio rebalancing effects. In addition, the learning process of investors may play a role. Meanwhile, the value effect, i.e. value stocks being more exposed to monetary policy risk than growth stocks, becomes more evident when double-sorted size-value portfolios are used for the estimations. Within each size quintile, the most value stocks are more responsive to changes in monetary policy conditions than the most growth stocks. Overall, the results provide some support for the credit channel of monetary policy transmission. Finally, the findings do not provide convincing evidence of the strong monetary policy reaction to stock price developments at either market or stock portfolio levels.

Overall, the empirical findings reported in the thesis have important implications for policymakers at the Federal Reserve and other central banks, economists, investors in financial markets and the man in the street. With respect to policy making, it provides a useful analysis of the effects of both conventional and, at the zero lower bound, unconventional monetary policies on financial markets. Asset prices constitute a part of the transmission mechanism of monetary policy to broader economy, thus, it is crucial to understand whether policy decisions have any influence over this stage of the transmission. Furthermore, unconventional policies were not ever used before to such an extent and their effects were not known prior to the implementation. Hence, the results reported here are important for future policy making at the zero lower bound. Financial market participants may find it valuable to have a better insight into what determines monetary policy decisions. This allows them to be able to anticipate future policy changes. Also, the thesis sheds some light on what may be the consequences of those changes on their investments. Finally, the general public could benefit from the analysis here as it helps to gain a broad understanding of how monetary policy is conducted and what impact it may have on financial markets and the economy, thus, it enables households to make better informed decisions regarding their finances.

There are several potential routes for future research following from this doctoral thesis. Firstly, Chapter 1 estimates forward-looking augmented Taylor rules until the zero lower bound period started. It may be of interest to estimate an interest rate-based policy rule for the Fed over the extended sample period and accounting for the zero lower bound and unconventional policies. For instance, one could estimate the Taylor rule using an interest rate-based measure of monetary policy stance that incorporates unconventional policy actions, such as the short-term shadow federal funds rate calculated by Wu and Xia (2016). Secondly, Chapter 2 examines solely government bond returns; however, the analysis could be extended to other types of bonds. Together with my co-authors I have already addressed this point by developing a working paper that applies a similar methodology as in Chapter 2 to the US corporate bond market. It is available online in the discussion paper series of the Adam Smith Business School, University of Glasgow. In addition, the analysis could possibly be adapted to also examine the impact of the Fed's forward guidance. Finally, Chapter 3 leaves out the analysis of unconventional monetary policy effects on stocks. Thus, the framework could possibly be amended to take into account the Fed's balance sheet policies during the financial crisis and its aftermath.

Finally, I would like to highlight that three working papers in total have been developed based on Chapter 1 and Chapter 2. All three papers are available online in the discussion paper series of the Adam Smith Business School, University of Glasgow.