# Essays in Contemporary Monetary Policy: Structural Change, Institutional Reforms and Interdependence

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## ABSTRACT

The four essays within this study investigate a series of issues recently emerged in monetary theory and policy. While the common theme is the empirical evaluation of the effects of structural change and institutional reforms, the perspective from which I study this problem markedly varies across the chapters.

The first two chapters of the thesis are more closely related. In the first chapter, I derive and estimate interest rate reaction functions for the G-3 economies and four inflation-targeting countries, to assess whether policy behaviour in these economies has significantly changed in recent years. Contrary to a commonly stated view, some interesting differences emerge amongst the policy rules followed by the central banks in the G-3 economies. Furthermore, the adoption of inflation targets and the move to greater central bank independence in the second group of countries do not seem to have significantly altered the way in which monetary authorities react to final policy objectives.

In the second chapter, I apply the same econometric methodology to an optimal interest rate rule derived for four former ERM -now EMU- countries. The existence of the exchange rate constraint permits to draw, *inter alia*, some empirically testable hypotheses about the effects of fiscal policy credibility on interest rate determination. My findings show that some of the economies faced, on their road to EMU, remarkable costs in achieving nominal convergence with Germany, mainly due to the presence of concerns on the sustainability of their fiscal stances.

The third chapter conducts an empirical investigation on the leading indicator properties of broad monetary aggregates for future inflation in the euro area. Using aggregate data for the area, I first test for Granger non-causality of M3 on prices, and then estimate a series of forecasting equations for inflation. My findings suggest that the information content of monetary aggregates, but also of the term spread and the output gap, is helpful for forecasting the behaviour of future inflation in the area. To this purpose, however, the joint use of information obtained from monetary models as the one adopted in this exercise, and from other, more "structural" models of the euro area, appears a superior forecasting strategy.

Finally, chapter four adopts a time-varying VAR perspective to obtain a tentative attribution of observed fluctuations of the bilateral real exchange rates between the USA, the UK and Italy, to shocks in relative productivity levels and the fiscal position. A Kalman filter approach is employed to assess the changing contributions of each of these variables, and of shocks to the monetary stance, to the behaviour of real exchange rates over the last 130 years. While confirming the relevance of fiscal shocks in triggering observed deviations of the exchange rate from PPP, the study finds little evidence in favour of persistent productivity effects on the real exchange rate. Moreover, the exchange rate regime in place plays a substantial role in determining how shocks are transmitted to the real exchange rate.

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## 1. Motivations

In recent years, the role played by policy institutions has become a centrepiece of macroeconomic theory. The literature on the political economy of monetary policymaking has repeatedly argued that the strategic interactions between markets and institutions strongly affect economic performance and social welfare. Furthermore, changes in the institutional settings of monetary policy, like the adoption of fixed exchange rate mechanisms and inflation target regimes, and the move towards greater central bank independence, have been extensively discussed by numerous theoretical contributions.

A first aim of this thesis is to bridge the substantial gap existing between some of the theoretical predictions elaborated during this intense debate, and the apparent lack of empirical validation that has often accompanied their emergence as normative guidelines for policymakers.

For instance, I shall derive and then estimate simple monetary policy rules for three groups of central banks: the G-3 economies, four countries that have recently adopted inflation targets, and four former ERM economies, now integrated into EMU. For the first group, I shall assess whether the observed conduct of monetary policies can be defined according to a recently emerged paradigm, whereby short-term interest rates have been long set to stabilise inflation expectations, with little, or no, concern for output stabilisation. For the second group of economies, I will evaluate whether the adoption of inflation targets and the move to greater central bank independence have significantly altered the way in which monetary authorities pursue final policy objectives. Finally, I will conduct a similar exercise on four EMU countries, to investigate the extent to which the existence of external constraints, like the ERM and the well-known fiscal convergence criteria, have affected central banks pursuit of domestic policy objectives.

The process of monetary unification in Europe, recently completed with the establishment of the European System of Central Banks, appears to raise a number of interesting questions. I shall devote the last chapter of this thesis to understanding, from an innovative perspective, the extent to which real and monetary shocks determine fluctuations in the real exchange rate. If real shocks turn out to have persistent effects on the real exchange rate, the relative desirability of fixed exchange rate systems and monetary unions, and of the recent proposal of dollarisation for some developing and transition economies, would have to be considered under a different, more critical perspective.

Finally, the third chapter will investigate the extent to which information about the behaviour of monetary aggregates can be helpful in predicting future inflation in the euro area. In addition, I shall evaluate the relative ability of monetary and "structural" models of euro area inflation in predicting future price developments.

Overall, I believe that this empirical research successfully addresses a number of relevant methodological issues, and unveils some relevant features concerning the role of structural and institutional change in modern macroeconomics. What follows is a brief description of the structure of the thesis, and of the way in which the single chapters will investigate the issues just sketched.

## 2. Thesis Structure

*First Chapter: Interest Rate Rules and Policy Shifts in OECD Economies*

In recent years, the theoretical literature on credibility, central bank independence, and monetary policy rules has greatly developed. Some genuine effort has been produced towards a deeper understanding, *inter alia*, of the way in which policy institutional settings affect economic performance. In addition, many contributions have stressed that the effectiveness of monetary policies depends on the way in which interactions between policy (and political) institutions and society as a whole are shaped.

The first chapter of the thesis aims at contributing to this (so far) chiefly theoretical debate, by providing some empirical evidence about the past effects of monetary reforms on the conduct of monetary policy. The goal is to study the consequences that institutional changes like the introduction of inflation targets, or the granting of a more independent status to the central bank, have had on the way in which policymakers react to the state of the economy. More in detail, we assume that the historical conduct of monetary policy in a number of countries can be effectively summarised by simple policy rules, which can be generated as a result of conventional optimising frameworks. It is further assumed that the behaviour of central banks in several OECD countries can be described according to a simple relationship between the policy instrument -usually a call money rate- and expected inflation and a measure of the business cycle. We thus check whether the reaction of monetary authorities to the state of the economy, as exemplified by such simple relationship, has radically changed during the past two decades.

We carry out such investigation by estimating interest rate reaction functions over the period 1970-1997, for the G-3 countries, and over 1980-1997 for a group of economies that have recently adopted an inflation-targeting framework (New Zealand, Canada, United Kingdom, and Sweden). We subsequently assess the stability of estimated equations and parameters, having two, closely related, aims. For the former group of countries, we want to check whether interest rate policies have undergone significant shifts in the past years. In the latter group, whether the introduced policy reforms have had any consequence on central banks' attitude towards inflation and cyclical conditions.

The above rules are usually derived in a context in which the central bank is assumed to face a given trade-off between inflation and output variability. It is then natural to think of a significant shift in the weight the rule assigns to, say, inflation as opposed to output stabilisation, as a change in the policymakers' –or the public's- relative preferences towards such final objectives. Alternatively, one can imagine such change as generated by some relevant institutional reforms, like the ones above mentioned, which may have significantly altered the emphasis on inflation control. In either case, structural instability displayed by the estimated equation and parameters can be overall interpreted as a signal of underlying structural change. Furthermore, the behaviour of estimated parameters over time can illustrate the extent to which the relative emphasis placed on alternative final objectives has changed during the period under investigation.

The horizons, over which policy rules as ours are estimated, are usually long. They span periods in which overall economic change makes models with time-invariant policy rules and macroeconomic structures not particularly robust. This is why we employ an appropriate Kalman filter technique and the Structural Time Series Approach to generate the regressors we use in the recursive estimation of our policy rules. That is, we assume that agents have limited information about macroeconomic variables. In particular, we hypothesise that the central bank and the private sector formulate their expectations about future inflation and output using only information available up to the time in which such expectations are formed. This way, our model allows for a fairly simple but essential learning process as regards inflation expectations, and for a flexible but not restrictive way of characterising central bank's information about the economy.

The key findings of our exercise can be summarised into two points.

First, although interest rate reaction functions for the G3 appeared overall stable over most of the sample we study, significant differences in the estimated parameters showed up when the models were re-estimated over a shorter period. In addition, contrary to the view according to which an implicit inflation targeting "attitude" can be found in all G-3 countries' observed policy conduct, we found evidence of a much more differentiated picture. Overall, it is only since the 1990s that estimated interest rate rules in these countries begin to look like the ones some research on inflation-forecast targeting has recently illustrated.

Furthermore, we found very little evidence supporting the view that central bank independence and the adoption of inflation targets have a substantial impact on central banks' conduct. While some signs of structural instability and parameter shift were easily detected in our estimated reaction functions, the timing of such changes did not always coincide with the announcement and/or the introduction of the reforms. We interpret this as a sign that institutional reforms, in the historical contexts we study, were brought in simply to consolidate gains in terms of lower inflation, or simply reflected a possible earlier shift in collective preferences towards the relative costs of inflation.

## *Second Chapter: Interest Rate Rules and Policy Credibility from ERM to EMU*

In the same vein as the previous chapter, here we wish to evaluate the path followed by monetary policies throughout Europe in the process of monetary convergence towards EMU. More in detail, we ask ourselves how the nominal convergence achieved amongst the former EMS countries was affected by the presence of the Exchange Rate Mechanism. In addition, we attempt a broad assessment of the way in which the required convergence in budget positions across countries has affected the response of national monetary authorities to final domestic objectives.

We thus estimate and evaluate interest rate reaction functions -similarly to what we did for the G-3 and inflation-targeting economies- for four key European countries: France, Italy, Ireland and Belgium. Our estimation sample -1980Q1-1997Q2- covers the whole period spanned by the EMS. It is then interesting to understand how the convergence of national Central Banks' -or the public's- preferences towards Bundesbank's anti-inflationary attitude took place. Similarly, given the relevance attributed by the Maastricht Treaty and the subsequent Stability and Growth Pact, to fiscal consolidation, we wish to evaluate how such additional constraint affected Member States' macroeconomic conditions on the road to EMU.

We provide some empirical evidence supporting the view that the historical path followed by monetary policies in the former Members of EMS to achieve nominal convergence with Germany was far from uniform. We show that the process itself bore significant shifts to the way in which monetary policy authorities reacted to domestic objectives. In addition, we allow our theoretical model to take into account the possibility that the credibility of the fiscal stance explicitly affected interest rate policies adopted in the EMS countries. With imperfect credibility, an unsustainable fiscal position in principle may induce markets to believe that the central bank will need to loosen its anti-inflationary stance (and, ultimately, the country's exchange rate commitment) in the future.

If unbalanced fiscal policies were to affect market perception of the probability of loose interest rate policies in the future, the optimal policy rule would directly target such perception. By consequence, and with reference to the policy rules analysed in the preceding chapter, the policy instrument would be explicitly reacting to a measure of the exchange rate risk, other than to final output and inflation objectives. The significance of the coefficient attached to this measure, proxied by the adjusted spread between long-term interest rates, in an estimated interest rate reaction function, would then signal the extent to which inflation and output stabilisation were sacrificed in the attempt to stabilise the exchange rate within the ERM band. In addition, the evolution of the way in which the central banks were reacting to the spread and to other regressors, and an assessment of the stability of estimated reaction functions, would illustrate further aspects. For example, it would show the extent to which the adoption of a tougher exchange rate commitment since late eighties (the "hard ERM"), and the varying commitment of national authorities to programs of fiscal consolidation, affected short-term interest rate determination.

Estimated interest rate reaction functions for the countries in our sample show that budget policies had severe effects in further constraining the behaviour of monetary authorities. In all countries, monetary policy stances seem to have been often motivated by the need to respond to changes in the credibility of the country's exchange rate position within the ERM band. Interestingly, during the "hard ERM" phase, in France, Belgium and Ireland, the importance of the long yield spread tends to decrease as severe efforts of fiscal retrenchment were undertaken. In such countries, the ERM turbulence in 1992-93 does not appear to have significantly affected interest rate policies, probably thanks to the largely achieved macroeconomic stability. On the contrary, for Italy, well-founded concerns surrounding its macroeconomic policies at the eve of Stage Three of EMU, appear to have severely constrained interest rate determination.

*Third Chapter: Assessing the Information Content of Euro Area Monetary Aggregates*

Economic theory suggests that money can play two roles in a monetary policy strategy. In standard inflation-forecast and monetary targeting regimes, the behaviour of monetary aggregates can be usefully monitored by the central bank to obtain information about future inflation. Monetary authorities adopting money growth as an information variable assume that past and current *monetary* developments contain useful information about current and future *price* developments.

On the other hand, in standard regimes of monetary targeting, the money stock is seen to provide for a *nominal anchor* to the whole economic system, and to inflation expectations in particular. Consistently with the view that inflation is, ultimately, a monetary phenomenon, the announcement of a target for the growth of some broad monetary aggregate helps the private sector forming expectations about future nominal variables.

It is thus clear that, for a monetary policy strategy aiming at using monetary aggregates, either as an intermediate target (nominal anchor role) or as an information variable (*leading indicator* role), the identification of the statistical properties of the money-prices relationships is critical.

Well before the start of Stage Three of the European Monetary Union, a number of empirical contributions have then addressed some of the above issues, from a number of perspectives. Using aggregate data for the euro area over the period 1980-1998, we devote the first part of this chapter to conduct a series of tests on the null hypothesis that money does not Granger-cause prices. The study is carried out within the context of a cointegrated VAR system. Subsequently, the leading indicator properties of M3 are investigated within the so-called *Pstar* framework. In other words, we try to assess the information content, and more specifically the predictive power, of the real money gap (actual minus equilibrium real money balances), but also of the output gap and a measure of the term spread, for future inflation. Finally, we perform a number of forecast encompassing tests, aimed at comparing the predictive ability of our forecasting model *vis-à-vis* that of a model in which developments in the money markets do not play any role in forecasting inflation.

The main conclusions of our analysis are as follows. First, there is very little empirical evidence against the null of Granger non-causality of M3 aggregates on prices for the euro area. Second, our forecasting model shows the existence of a reasonably strong positive association between the real money gap and future inflation up to five-six quarters ahead. However, the output gap and, to a lesser extent, the real interest rate or the term spread, tend to display similar predictive ability. This is in line with some influential recent findings for the US, and allows concluding that standard *Pstar* models are likely to forgo the rich information content of variables other than the real money gap. Finally, detailed forecast encompassing tests suggest that information from "monetary" models like ours and that from more "structural" models of the euro area should be systematically and jointly used to study future price developments.

*Fourth Chapter: Time-Varying Perspectives on Real Exchange Rates, Productivity Levels and Government Spending*

An extensive body of contributions has clarified that exchange rates do tend to deviate from Purchasing Power Parity (PPP). Although a more decisive disentangling of the reasons of such behaviour will need to resort to further advances in econometric theory, the main findings of the recent empirical literature on real exchange rates can be summarised as follows. First, PPP does appear to hold, but only in the very long run: both in its absolute and relative versions, PPP fails to hold. Second, the observed departures of the exchange rate from PPP are more persistent than traditional, flexible-price models of the exchange rate would predict.

This chapter aims at providing some empirical evidence as to the *origins* of the observed deviations of the exchange rate from PPP. The question of whether PPP holds, and the related one about the mean-reverting properties of the nominal exchange rate, yet deserve exhaustive empirical work, but our focus here is different. With our study, we wish to shed some light on the *nature* of the movements of the real exchange rate.

Thus far, the existing empirical evidence has provided limited support to the idea that sustained divergences in government spending and sectoral productivity patterns might be at the root of persistent fluctuations of bilateral real exchange rates. Moreover, the persistence of deviations of the nominal exchange rate from PPP appears to critically depend on the exchange rate regime in place. That is, during periods of floating exchange rates, persistent fluctuations of the real exchange rate seem to be more common than under fixed exchange rates. Also, market distortions -in the form of pricing-to-market behaviour or similar market segmentation practices- may prevent perfect goods arbitrage, and make the dynamics of relative prices diverging from that of nominal exchange rates.

Against this background, we then seek to attribute the observed movements of the real exchange rates between the USA, UK, and Italy, to some of the causes above summarised. That is, we try a tentative attribution of the shocks to the real exchange rates to existing divergences in the fiscal stance, differential productivity levels, and significant differences in monetary conditions.

There are two main novelties in our study. First, following some recent advances in Bayesian approaches to the estimation of Vector Autoregressions, we employ a time-varying methodology. We apply such techniques to estimate an unrestricted VAR comprising the bilateral real exchange rate, a measure of productivity differentials, an indicator of the relative fiscal stance, and the ratio between real *ex-post* interest rates in the two countries. We then decompose the total residual variance of each VAR equation into stochastic contributions attributable to innovations in each endogenous variable. By applying a Kalman filter technique to the system's estimated parameters and variances, we decompose the total variance of the real exchange rate equation into contributions that are allowed to change over time. The particular state-space representation we adopt for our models enables us to impose the least restrictive assumptions on the dynamic structure of the system; VAR coefficients are simply assumed to be stationary white noise processes. This way, our model is able to pick up the changing influences of the monetary regime, of productivity shocks, and the fiscal and monetary stance, on the real exchange rate, over time. We further avoid imposing any particularly stringent structure for the identification of the shocks, by using Generalised Impulse Responses to examine the dynamic response of the real exchange rate to shocks in each of the remaining variables of the system.

The second novelty of this study is represented by the data we use. We apply our approach to a long historical set of annual observations spanning the last 130 years and obtained from a variety of sources. The bilateral real exchange rates between the USA, UK, and Italy are analysed across all international monetary regimes (classical Gold Standard, Bretton Woods, the post-1973 floating and the EMS) these countries were involved in during the sample period.

The main results we obtain can be summarised as follows. First, we find very little evidence in favour of shocks to relative productivity levels as having persistent appreciating effects on real exchange rates. Such result, though not valid across all the sub-sample estimates we obtain, appears to be overall quite robust. Second, we find some stronger evidence in favour of shocks to the relative fiscal position of a country as triggering substantial real exchange rate fluctuations. Third, estimates conducted within single exchange rate regimes suggest that the response of real exchange rates to shocks in the other variables appears to critically depend not just on the international monetary arrangements historically in place, but also on demographic and technological factors.

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Materials from **Chapter Two** are going to be used in the development of another paper: A. Muscatelli, P. Tirelli, and C. Trecroci, "Monetary Policy on the road to EMU: the dominance of external constraints on domestic objectives". University of Glasgow, 2000. I am sincerely grateful to my co-authors for their precious co-operation in both projects.

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The usual disclaimer applies.

# Forward-Looking Interest Rate Rules and Monetary Policy Regime Changes in OECD Economies

# Chapter 1

*"Every model is an oversimplification. Economies do change over time. Econometric equations often fail subsample stability tests. Econometric problems like simultaneity, common trends, and omitted variables are ubiquitous in nonexperimental data. The Lucas critique warns us that some parameters may change when policy does. Yet what are we to do about these problems? Be skeptical? Of course. Use several methods and models instead of just one? Certainly. But abandon all econometric modelling? I think not. The criticisms of macroeconometrics are not wrong, but their importance is often exaggerated and their implications misunderstood. These criticisms should be taken as warnings --as calls for caution, humility, and flexibility of mind- not as excuses to retreat into econometric nihilism. It is foolish to make the best the enemy of the moderately useful."*

Blinder (1998), p. 8.

## 1. Introduction

In recent years, the theoretical literature on credibility, central bank independence, and monetary policy rules has enormously expanded. Some common scepticism -not always fully justified- about the stabilisation effects of fiscal policies, has often translated into a genuine effort towards a deeper understanding, *inter alia*, of the way the institutional setting of monetary policy affects economic performance. In addition, such aim has often coincided with the need to clarify the general idea that the effectiveness of economic policies depends on the way interactions between policy (and political) institutions and society as a whole are shaped.

Some of the findings of this burgeoning research have greatly contributed to an improved knowledge of both the normative and positive aspects of policy design. In an era of important monetary reforms, like the establishment of a single monetary authority in Europe, the benefits of such efforts are invaluable, and eminently practical[1]. After having stressed the importance of price stability as the dominant goal of monetary policy, and of central bank independence as a guarantee against sub-optimal policy outcomes (Persson and Tabellini, 1999), the attention of the literature (and of central bankers) has recently re-focused on optimal policy rules. Remarkable effort has been devoted to envisaging the optimal solution to the classical policy problem: choosing the correct instruments and time

[1] Although some scholars' apparent dissatisfaction with the way policy design is actually implemented appears emblematic. See Svensson (2000b), for example.

paths to maximise the assumed policy objective function, subject to constraints on the economy's behaviour and policy institutional setting[2].

The present study aims at contributing to this (so far) chiefly theoretical debate, by providing some empirical evidence about the effects of monetary reforms on the conduct of monetary policy. Our goal is to study the consequences that institutional changes like the introduction of inflation targets, or the granting of a more independent status to the central bank, have had on the way in which policymakers react to the state of the economy. More in detail, we assume that the historical conduct of monetary policy in a number of countries can be effectively summarised by simple policy rules[3], which can be generated as a result of more or less articulated optimising frameworks[4]. We further assume that the behaviour of central banks in several OECD countries can be described according to a simple relationship between the policy instrument -usually an overnight interest rate- and expected inflation and a measure of the business cycle (Clarida, Gali, and Gertler, 1998). We thus check whether the way monetary authorities react to the state of the economy, as exemplified by such simple relationship, has radically changed during the past two decades. In addition, we seek to test whether major events in some countries' recent monetary history, like the introduction of inflation targets and central bank independence, have significantly affected authorities' behaviour. We carry out such investigation by estimating interest rate reaction functions over the period 1970-1997, for the G-3 countries, and for a group of economies that have recently adopted an inflation-targeting framework (New Zealand, Canada, United Kingdom, and Sweden). We subsequently assess the stability of estimated equations and parameters, having two, closely related, aims. In the former group of countries, we want to check whether interest rate policies have undergone significant shifts in the past years. In the latter group, whether the introduced policy reforms have had any consequence on central banks' attitude towards inflation control and cyclical conditions.

Interest rate policies are in practice complex decisions, relying on a multitude of indicators and models (Bank of England, 1999; Vickers, 1999), and by nature related to events not always captured by relatively simple econometric models. Nonetheless, there is now a wide consensus (Amato and Laubach, 1999; Batini and Haldane, 1999; Peersman and Smets, 1999; Gerlach and Schnabel, 1999) on the fact that the class of simple policy

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[2] Taylor (1999) collects a significant number of the most influential contributions in the area. See also McCallum (1999).
[3] The prime reference here is Taylor (1993).
[4] See, again, the essays contained in Taylor (1999), and Svensson (1997b)

rules examined here generates stabilisation properties very close to those displayed by optimal feedback rules. It is then reasonable to think that stability analysis conducted on estimated forward-looking interest rate rules as the ones we examine, can provide some indications about changes that eventually took place in the actual conduct of monetary policies.

More precisely, the above rules are usually derived in a context in which the central bank is assumed to face a given trade-off between inflation and output variability. It is then natural to think of a significant shift in the weight the rule assigns to, say, inflation as opposed to output stabilisation, as to a change in the policymakers' –or the public's- relative preferences towards those final objectives. Alternatively, one can imagine such instabilities as generated by some relevant institutional reforms, like the ones above mentioned, which have significantly altered the emphasis on inflation control. In either case, structural instability displayed by the estimated equation and parameters can be overall interpreted as a sign of underlying structural change. Furthermore, the behaviour of estimated parameters over time could illustrate the extent to which the relative emphasis placed on alternative final objectives has changed during the period under investigation. Although the latter task is better undertaken by explicitly assuming time-varying DGPs for the model's parameters -as we do, in a very different context, in the final chapter- some interesting evidence could be provided even in a less complex context.

There is, however, a serious element of caution to be borne. The horizons over which policy rules like ours are estimated, are usually long. They span periods in which overall economic change makes models with time-invariant policy rules and macroeconomic structures not particularly robust. Aside from model and parameter uncertainty, and considerations related to Lucas' critique, one of the dangers of such empirical exercises is to forget that alternative exchange rate regimes, financial innovation, and changes in the underlying structure of the economy, have fundamental influences on policy rules. This is why we employ an appropriate Kalman filter technique and the Structural Time Series Approach (Harvey, 1989; Kim and Nelson, 1999) to generate the regressors we use in the recursive estimation of our policy rules. That is, we assume that agents have imperfect information about economic variables. In particular, we hypothesise that the central bank and the private sector formulate their expectations about future inflation and output using only information available up to the time in which such expectations are formed. The result is that our model allows for a fairly simple but essential learning process as regards inflation expectations, and for a flexible, unrestrictive way of characterising central bank's information about the economy. This is probably why our estimates, while broadly confirming findings from previous studies, yield some innovative results as regards parameter estimates. In addition, results from sub-sample estimation of the same model, and structural stability analysis, tend to confirm that the properties of estimated policy rules should be carefully evaluated. That is, substantial caution should be adopted when observed policy behaviour is employed as a benchmark for optimal policy design.

The key findings of our exercise can be summarised into two points.

First, although interest rate reaction functions for the G3 appeared overall stable over most of the sample we study, significant differences in the estimated parameters showed up when the models were re-estimated over a shorter period. This was particularly evident in the US case. In addition, contrary to the view (Clarida, Gali and Gertler, 1998, for example) wherby an implicit inflation targeting "attitude" can be found in all G-3 countries' observed policy conduct, we found evidence of a much more differentiated picture. Overall, it is only since the 1990s that estimated interest rate rules in these countries begin to look like the ones some research on the inflation-forecast targeting approach has recently illustrated.

Furthermore, we found very little evidence supporting the view that central bank independence and the adoption of inflation targets have substantial impact on central banks' conduct. While some signs of structural instability and parameter shift were easily detected in our estimated reaction functions, the timing of such changes did not always coincide with the announcement and/or the introduction of such reforms. Estimated interest rate policy rules in Canada and New Zealand, and to a lesser extent in the United Kingdom and Sweden, did not display substantial instability in correspondence of such changes. We interpret this as a sign that those institutional changes, in the historical contexts we study, were brought in simply to consolidate gains in terms of lower inflation, or simply reflected a possible prior shift in collective preferences towards the relative costs of inflation.

The chapter is organised as follows. **Section 2** lays out the simple benchmark framework used to derive the optimal policy rule we subsequently estimate. With the help of such benchmark model, we examine in **Section 3** some of the major issues in the recent literature on optimal policy rules. In **Section 4** we illustrate our results for the US, Japan and Germany, highlighting important and, so far, downplayed, differences amongst monetary policies in these countries. **Section 5** turns to the inflation targeting experiences of New Zealand, Canada, United Kingdom and Sweden. In **Section 6** we draw some conclusions.

## 2. Forward-Looking Interest Rate Rules in a Simple Benchmark Model

As we mentioned in the Introduction, the search for an optimal solution to the policy problem has produced an impressive amount of studies. As exhaustive surveys of this huge literature are now readily available (Persson and Tabellini, 1999; Clarida, Gali and Gertler, 1999; Christiano and Gust, 1999)[5], we avoid getting through all single aspects of the issue, and focus on major points, with the help of a simplified theoretical model.

The key point of all recent analyses of monetary policy is the assumption that some nominal frictions allow interest rate changes to affect real variables in the short run. The nominal rigidities postulated by the most recent optimising models for the evaluation of monetary policy vary considerably. However, the two main strands attribute such frictions to stickiness in price-setting behaviour (Rotemberg and Woodford, 1999), or to rigidities in the money market (Christiano, Eichenbaum and Evans, 1997).

Interest rate policy is forced to take adequate account of forecasts about chosen final objectives such as inflation and output. The most intuitive reason for this is that policy actions affect the final goals only with some lag. First, it is now universally accepted that the response of real variables, like output and employment, to a monetary impulse materialises with a substantial lag. Second, changes in such variables in turn will influence the price level after additional time.

However, these are crucial, but still somehow traditional (Blanchard, 1990; Walsh, 1998) aspects of monetary policy evaluation. The real innovation of the recent wave of inter-temporal optimising models of interest rate determination is that the behaviour of private agents explicitly depends on what they expect about the future course of monetary policy. In other words, private sector's belief about how monetary policy operates affects the credibility of the policy stance, and determines the extent to which real

[5] See also the above mentioned contributions in Taylor (1999).

and nominal variables will respond to policy impulses. Any technology that enhances policy credibility will make inflation and/or output stabilisation closer to the outcome that one would obtain in absence of market imperfections.

The latter principle was already clearly established in the pioneering analyses of Kydland and Prescott (1977), and Barro and Gordon (1983). What instead recent forward-looking policy models have made clear (Svensson, 1997b) is that, given the existence of the above mentioned lags, and the forward-looking nature of private sector's behaviour, it is optimal for monetary authorities to set policy decisions accordingly. In other words, the policy instrument -usually a short-term interest rate- must be set in a way that the forecast of the objective variable, conditional on all available information and assuming unchanged interest rates, coincides with the policymakers' target for that variable. Among the first to model the credibility problem of monetary policy in the explicitly forward-looking fashion we have just sketched, is Svensson (1997b; 2000a). It is then with reference to this theoretical set-up that we shall illustrate our simple model.

In recent years, a widely accepted macroeconomic framework has become the basis for the majority of monetary policy studies. A relatively broad consensus has emerged about the fact that traditional Keynesian macromodels, although extremely powerful in generating key findings, were not sufficiently microfounded. A new wave of models, in which aggregate relationships were explicitly derived from the optimal behaviour of households and firms, thus began to be developed (Walsh, 1998; Rotemberg and Woodford, 1999; Bernanke, Gertler and Gilchrist, 1999).

The resulting behavioural relationships allow current aggregate values for macroeconomic variables to depend, *inter alia*, on the future course of monetary policies.

For example, let $y_t$ and $y_t^*$ be the current and potential level of output[6], and $p_t$ the price level. The following expectations-augmented Phillips curve can then be derived[7] from the basic, two-period, household's optimal consumption problem:

$$\pi_t = p_t - p_{t-1} = \lambda \pi_{t+1}^e + \varphi\left(y_t - y_t^*\right) \qquad [1]$$

[6] We might have equivalently assumed that $y_t^*$ represents the natural or NAIRU level of output.
[7] See Bernake, Gertler and Gilchrist (1999) or Rotemberg and Woodford (1999) for the analytic underpinnings of all aggregate relationships, which we do not illustrate.

In [1], current inflation depends on the inflation expected in the next period, and on current output gap[8]. In turn, the latter is affected by the deviations of the nominal interest rate from its expected value, and by a white noise shock:

$$y_t - y_t^* = -\gamma\left(R_t - R_t^e\right) + \varepsilon_t \quad [2]$$

Equation [2] is a standard relationship in the literature (Svensson, 1997b, 2000a). In line with a simple aggregate demand-aggregate supply framework, it postulates that a positive surprise in the interest rate level negatively affects current output. The Fisher *ex ante* parity holds, so that

$$R_t^e = r_t^* + \pi_{t+1}^e, \quad [3]$$

where $r^*$ is the real interest rate.

Next, we assume that monetary policy objectives involve stabilising, in each period, deviations of current inflation and output from respective targets $\tilde{\pi}$ and $\tilde{y}$. One crucial feature of these models is that the authorities' target level of output is assumed to lie above the natural level (Barro and Gordon, 1983; Cukierman, 1992; Rogoff, 1985; Svensson, 1997a). The central bank then has a short-run incentive to set its policy instrument below the level expected by the private sector, in the attempt to push output above its market-clearing level. The private sector, in turn, perceives the existence of such incentive, and adjusts its expectations accordingly. The classical result is the emergence of an inflation bias, which can be partially avoided only in presence of a technology that credibly constrains the future policy course.

It should be noted, however, that the existence of an overly ambitious output objective of the central bank is not strictly vital for a credibility problem to crop up. Substantial gains from commitment to a credible rule can emerge even in the absence of such distortionary preferences on the part of monetary authorities, simply because of the forward-looking behaviour of private agents. A clear and credible commitment of the monetary authority on inflation control makes the overall short-run output/inflation trade-

[8] The early Calvo (1983) and Rotemberg (1983) models with costly price adjustment yield essentially the same implication.

off faced by monetary authorities more favourable, thus reducing the size of the inflation bias[9].

More in general, Rotemberg and Woodford (1999) show that objectives like those sketched above can be derived directly from the minimisation of society's welfare function.

To sum up, we assume with Svensson (2000a) that the central bank attempts to minimise interest rate changes, and the departures of the policy instrument from its expected value:

$$L_t = \chi\left(\pi_t - \bar{\pi}\right)^2 + \left(y_t - \bar{y}\right)^2 + \rho_1\left(R_t - E\{R_t\}\right)^2 + \rho_2\left(R_t - R_{t-1}\right)^2 \qquad [4]$$

The above loss function reflects the fact that, since interest rate changes are a costly instrument of stabilisation policy[10], shocks are never fully stabilised in the longer term. Svensson (1997a) formalises this principle by assuming that authorities penalise deviations of the instrument from zero. In our model, instead, we do hypothesise that the policymaker knows the level of the interest rate consistent with the current state of expectations, but that he/she decides whether or not to deviate from it whenever some shock hits the economy.

In absence of any institutional device aimed at enhancing credibility[11], the model is solved by minimising [4] with respect to the policy instrument $R_t$, subject to [1] and [2]. This yields:

$$R_t = \Lambda r^* - \delta_0 + \delta_1 \pi^e_{t+1} + \delta_2 \varepsilon_t + \delta_3 R_{t-1} \qquad [5]$$

where the coefficients are combinations of model parameters as follows:

[9] For more details on this tangential issue, see Clarida, Gali and Gertler (1999), section 4.

[10] See Goodhart (1996) for some justification. See also Goodhart (1999) for further discussions.

[11] Friedman and Kuttner (1996) forcefully argue that the substantial inflation fall of recent years has taken place in the absence of such arrangements in all major OECD economies.

$$\Lambda = \frac{s^2\varphi^2\chi + s^2}{\left[s^2\varphi^2\chi + s^2 + \rho_2\right]};$$
$$\delta_0 = \frac{s\left(\tilde{y} - y^*\right) + \chi\varphi s\pi^*}{\left[s^2\varphi^2\chi + s^2 + \rho_2\right]}; \quad \delta_1 = \frac{s^2\varphi^2\chi + s^2 + \chi\beta\varphi s}{\left[s^2\varphi^2\chi + s^2 + \rho_2\right]}; \qquad [6]$$
$$\delta_2 = \frac{s\varphi^2\chi + s}{\left[s^2\varphi^2\chi + s^2 + \rho_1 + \rho_2\right]}; \quad \delta_3 = \frac{\rho_2}{\left[s^2\varphi^2\chi + s^2 + \rho_2\right]}$$

In line with Clarida, Gali and Gertler (1998), the system is stable when $\chi\beta\varphi s > \rho_2$, which in turn implies that $\delta_1 > 1$, and that the expected inflation response to the output gap be positive in equation [8] below. That is, the reaction function in [5] produces a stabilising response of interest rates to inflation movements, as long as the coefficient on expected inflation is larger than one. More intuitively, optimal policy entails adjusting the nominal interest rate to the extent that the involved change in the real rate fully offsets the observed variation in expected inflation. Such condition is fulfilled, as we have just seen, when the costs associated with interest rate changes, $\rho_2$, are not too big. It will be important to recall this critical condition when we evaluate the results of our estimation.

Similarly, central bank's concern for direct output stabilisation is captured by the coefficient associated with the supply shock, which is usually proxied by the deviations of actual output from its potential level.

Finally, the extent to which interest-rate-smoothing considerations affect policy decisions is picked up by the last term in [5], and it is directly related to the perceived interest rate adjustment costs.

Another important point to note is that the interpretation of [5] as a testable reaction function calls for some caution as to the meaning to attribute to the constant term $\Lambda - \delta_0 r^*$. In Clarida *et al.*'s (1998) reaction function, this is simply interpreted as the long-run component of the real interest rate. In our relationship, which is a very close analogue to what estimated in that study, the constant is a function of the real interest rate, but also of the inflation target and the inflation bias, as reflected by the difference between central bank's output target and the potential level. This certainly calls for some caution when it comes to evaluating estimated parameters derived from alternative assumptions about central bank's preferences.

Under rational expectations and full information on private sector's expectations, no systematic policy surprises are possible. That is

$$R_t^e = E_{t-1}\{R_t\}, \qquad [7]$$

which in turn allows us to calculate equilibrium *expected* inflation,

$$\pi_{t+1}^e = r^*\left[\frac{\rho_2}{\chi\beta\varphi s - \rho_2}\right] + \frac{s\left(\tilde{y} - y^*\right) + \chi\varphi s\pi^*}{\chi\beta\varphi s - \rho_2} - R_{t-1}\left[\frac{\rho_2}{\chi\beta\varphi s - \rho_2}\right] \qquad [8]$$

and equilibrium *actual* inflation:

$$\begin{aligned} \pi_t = r^*\left[\frac{\beta\rho_2}{\chi\beta\varphi s - \rho_2}\right] + \frac{\beta\left[s\left(\tilde{y} - y^*\right) + \chi\varphi s\pi^*\right]}{\chi\beta\varphi s - \rho_2} + \\ - R_{t-1}\left[\frac{\beta\rho_2}{\chi\beta\varphi s - \rho_2}\right] + \varphi\left(1 - s\delta_2\right)\varepsilon_t \end{aligned} \qquad [9]$$

In [8] and [9], one can clearly see that the classical Barro-Gordon inflation bias can in principle be offset by a suitable choice of the inflation target (Svensson, 1997a).

A further complication of this very basic picture, however, derives from the fact that so far we have assumed monetary authorities' preferences as fixed, and perfectly known to the private sector. Moreover, the authorities have full information about the state of the economy. In such unwarranted case, inflation and interest rates will be stationary stochastic processes. However, in what follows we allow for imperfect information as to central bank's objectives, and we also assume that the latter's ability to predict the supply shock is limited.

Following Muscatelli (1998, 1999), Faust and Svensson (2000), and Walsh (1998), we let private sector's beliefs about central banker's relative inflation aversion to be represented by the a simple updating process:

$$\chi_t^e = \phi\chi_{t-1} + \kappa_t \qquad \kappa_t \sim (0, \sigma_\kappa^2) \qquad [10]$$

In addition, suppose that the policymaker has imperfect knowledge of the state of the economy (the supply shock $\varepsilon_t$), that he/she makes inferences on it through a forecasting process, and that such forecast, $\varepsilon_t^f$, is private information. In each period, then, private agents are uncertain as to whether the shock they observe is due to a true supply shock, or it simply reflects a shift in policymaker's preferences $\kappa_t$ (Cukierman, 1992). Private sector's perception of the interest rate rule will thus be different from [5], and will be articulated as

$$R_t = r^* - \mu_0 + \mu_1 \pi_{t+1}^e + \mu_2 \varepsilon_t^f + \mu_3 R_{t-1} \qquad [11]$$

In [11], the parameters are linear functions of those in equation [5], but now $\chi_t^e$ has replaced $\chi_t$, and the supply shock is only the forecast of the one on the right-hand side of [5]. The standard signal-extraction problem faced by private agents then translates into assuming that agents update their expectations about the business cycle and central banker's preferences each period, by looking at past disturbances' variances.

This mechanism allows us to conclude that, in the case of a sudden regime change, like the introduction of some institutional device aimed at better controlling inflation -inflation targets or the granting of an independent status to the central bank- we have two possible scenarios. If the new regime is a fully credible one, the adjustment of equilibrium inflation and interest rates is immediate. If the reform is instead only partly credible, nominal variables will adjust gradually to the new steady state. Assuming one can obtain estimates for the regressors of the above relationship, significant and permanent changes of estimated coefficients in [11] could be easily detected. These, along with eventual instability of the overall equation in correspondence of major policy shifts, would signal either changes in policymakers' preferences, or the introduction of some institutional reforms, or both[12]. Clearly, the interpretation of observed shifts in our estimated parameters and functions has to be carefully conducted, but in general, major and permanent shifts in estimated coefficients can be attributed to parallel changes in policy preferences.

To sum up, what we aim to obtain, by studying estimated versions of reaction functions like [11], is an assessment of the stability of central bank's conduct. Of course, we do not assume that the latter is fully described by the simple rule above illustrated. Although the point is not always made clear in various contributions to the field, we acknowledge that in practice central banks respond to a variety of indicators. They do gather information about the current state of the economy using a host of economic and econometric models (Bank of England, 1999; Vickers, 1999). Nevertheless, we believe that estimated versions of the Taylor rule (Taylor, 1993; Gerlach and Smets, 1999) or of our relationship in [11], capture the way in which monetary policies translate in a simple rule expressed in terms of expected inflation and output gap. We can clarify the point with reference to a concrete example. As is universally known (von Hagen, 1995, 1999; Issing, 1997), Bundesbank's announcements of annual money growth targets since 1974, provided agents with a reliable corridor for expectations about the future policy course. Although such intermediate targets were in reality missed in more than 50% of the cases, this did not substantially impair inflation and output stabilisation. Consequently, finding a stable reaction function for Germany[13] would amount to vindicate the idea that the overall set of operational rules adopted by the Bundesbank was stable over time, and that the reliance on intermediate targets did not take place at the expense of achieving the final objectives. On one hand, this motivates our subsequent choice (and Clarida *et al.*'s, 1998) of including additional regressors to the baseline specification in [11].

A further possibility of instability in estimated reaction functions crops up for those experiences, like the US, the UK, and Canada, where the instability of monetary aggregates led to their early demise as an intermediate policy target. Equivalently, we argue that in countries (Japan, and again the UK, Canada and Sweden) where links with benchmark foreign exchange rates (or pegs) followed a similar fate, it will be likely to detect similar instabilities. This is why we shall test whether variables like lagged money growth, leading foreign interest rates or exchange rates, significantly enter the policy rule.

What really matters when one evaluates the effectiveness and stability of policy rules, is their performance in terms of announced final, not intermediate, objectives. Should any of the above indicators enter an estimated reaction function, one would then conclude that the role it plays in the policy rule is similar to those played by final inflation and output objectives. Again, should we really find, say, a significant role for M3 growth

[12] Of course, a third possibility would be that the observed instability is simply due to changes in the underlying behavioural relationships between the variables of interest. The way in which we calculate the series for expected inflation and the output gap, however, takes into account such possibility.

[13] This is the case with our estimates.

in the estimated interest rate function for Germany[14], we would conclude that the stabilisation of money supply in Germany took place at the expense of meeting final output and inflation objectives.

The key problem with the estimation of reaction functions like [11] is the availability of the unobserved series for expected inflation and potential output, along with the identification of some updating mechanisms for all expected variables. Below we show that one optimal, though not unique, way of solving this issue is applying the Kalman filter to those variables. As we will see in Section 5, some authors have followed alternative routes (Clarida *et al.*, 1998; Gerlach and Smets, 1999; Favero and Rovelli, 1999). Before turning to better explain the way in which we tackled the issue, we now proceed to a brief discussion of the recent literature on monetary policy rules, trying to isolate specific empirical aspects still to be investigated.

## 3. Estimating Interest Rate Reaction Functions: Existing Theory and Evidence, and the Way Forward

### *3.1 Forecast-Based Monetary Policy Rules*

The stabilising and welfare properties of interest rate rules, and of monetary policy rules more in general, have been extensively investigated in recent years. This surge of interest is certainly related to the experience of many countries that have chosen to employ inflation targets in their monetary policy strategy. As in the majority of OECD economies, in such countries the monetary authority is charged with achieving and maintaining price stability. In inflation targeting regimes only, the latter is explicitly defined in terms of a numerical objective for annual inflation. While there is no equivalent provision for an output target, it is widely believed that if the bank attempts to hit the announced inflation target on a period-by-period basis, the consequent instrument fluctuations would invariably involve substantial losses in terms of output variability (Goodhart, 1996, 1999).

Most theoretical contributions (Svensson 1997b; Amato and Laubach, 1999; Goodhart, 1999; Rotemberg and Woodford, 1999; Rudebusch and Svensson, 1999) have

[14] This was not the case with our estimates. Indeed, we conclude below that, in line with what found by Bernanke and Mihov (1997), and Clarida and Gertler (1997), Germany appears to be, at least on an implicit basis, an expected inflation targeting regime.

then made clear two points. First, the forward-looking nature of private agents' behaviour is such that welfare is in general maximised by stabilising the *forecast* of inflation around the appropriately chosen target, at some horizon. Second, central banks can rein in *actual* inflation by adjusting interest rates in order to stabilise those forecast: in so doing they can lessen output variability. In other words, interest rates must be set, in practice, to minimise the departures of forecast inflation at a specific horizon from the assumed target value.

The policy problem is thus partly relocated. What becomes now crucial is:

a) the choice of indicators used to formulate inflation forecasts;
b) the appropriate choice of the inflation target -whether a point, a corridor, an asymmetric range, and the price index on which it is defined; and
c) the horizon over which the forecast is stabilised around the target level.

The role of indicator variables for monetary policy forecasting is somehow studied in Chapter 3 (see also Gerlach and Svensson, 1999, and Svensson and Woodford, 2000). Moreover, the issues surrounding the choice of the optimal inflation target are not the focus of our present investigation (see IMF (1999) for a good survey).

The European Central Bank's monetary policy strategy explicitly argues that the objective of price stability is to be pursued over a medium horizon (ECB, 1999). This is a clear recognition of the fact that attempting to hit an inflation target on a period-by-period basis would involve significant output losses. Of course, ECB's (as well as US' and Japan's) policy strategy does not involve an explicit inflation-forecast approach à la Svensson (1997a, b)[15]. However, forward-looking rules like the ones derived in the previous section and discussed here are quasi-optimal tools, whose validity readily extends to whatever institutional setting is in place (Rudebusch and Svensson, 1999; Bernanke *et al.*, 1999; Batini and Haldane, 1999; Amato and Laubach, 1999; Faust and Svensson, 2000).

Following Batini and Haldane (1999) and Svensson (1997b), the most studied form of inflation-forecast interest rate rule is

$$r_t = \theta r_t^* + \varphi r_{t-k} + \gamma E_t \pi_{t+j} + \lambda \left( y_t - y_t^* \right), \qquad [12]$$

[15] See also Muscatelli (1999), Walsh (1998), and the critical remarks on ECB strategy contained in Dornbusch, Favero and Giavazzi, (1998) and Svensson (2000b).

where $r_t$ is the short-term ex-ante real interest rate, $r_t^*$ represents its long-run equilibrium value, and $E_t\pi_{t+j}$ indicates the $j$-period ahead inflation rate expected in time $t$.

The presence of lagged terms of the real interest rate accounts for the observed interest-rate smoothing behaviour of many central banks. It also allows for the possibility that interest rate changes might be postponed to avoid undesired disruption of the current level of economic activity (Batini and Haldane, 1999). The latter argument implies also that the presence of an output gap term in [12] might not be strictly necessary to account for authorities' attention towards output stabilisation. By tuning the degree of interest-rate smoothing and the lead in the inflation forecast, the bank can easily trade off *output* volatility for *inflation* volatility.

Eq. [12] can be trivially re-parameterised in terms of the nominal short-term interest rate:

$$R_t = \alpha + \varphi R_{t-k} + \eta E_t \pi_{t+j} + \lambda\left(y_t - y_t^*\right) \qquad [13]$$

where now $\eta = 1 + \gamma$, while the constant $\alpha$ includes both the long-run real interest rate and the persistence of the inflation forecast. The simple theoretical policy rule derived in Section 2,

$$R_t = r^* - \mu_0 + \mu_1 \pi_{t+1}^e + \mu_2 \varepsilon_t^f + \mu_3 R_{t-1} \qquad [14]$$

can then be amended to obtain something akin to [13]. In particular, by including a longer lead for expected inflation, a longer lag for the interest-rate smoothing term, and substituting the output gap for the forecast supply shock $\varepsilon^f$, one obtains a reaction function defined in terms of the policy instrument. Indeed, what we shall estimate is the following:

$$R_t = \mu_0 + \sum_{i=1}^{k} \mu_i R_{t-i} + \vartheta E_t \pi_{t+j} + \lambda\left(y_t - y_t^*\right) \qquad [15]$$

We estimated the above relation using simple Recursive Least Squares. This seemed to comply with the need of simplifying the statistical burden of our exercise, while still providing effective means of testing the various issues at stake. An obvious alternative would have been using Full Information Maximum Likelihood techniques[16]. We did run some explicit FIML estimation[17], but the results we obtained were not substantially different, in nature, from those below illustrated. Our results tend to show that a maximum length $k = 2$ is sufficient to pick up the extent to which all central banks included in our study smooth interest rate. After having estimated a baseline specification of [15], we searched for the appropriate lead $j$ for the inflation forecast, using conventional goodness-of-fit criteria. In other words, after estimating our baseline relationship with $j = 4$, we attempted with alternative lead lengths, selecting for our final specification the one best performing in terms of R-squared, regression standard errors, residuals' stability, etc.

The specification above then allows us to empirically evaluate many of the issues the literature on optimal forecast-based policy rules has introduced. The estimated weight the central bank places on the expected inflation, and the lead-length of such term ($\vartheta, j$), reveal: a) the degree of aggression with which the bank reacts to changes in the inflation forecast, and b) the extent of "forward-lookingness" of bank's behaviour. The parameters ($j, k, \varphi$) in turn capture the overall degree of inertia in interest rate policy. Finally, a significant value for $\lambda$ would reflect bank's concern over output, over and beyond the one showed, for the reasons earlier illustrated, by the combination of interest-rate smoothing terms and lead/lag parameters[18].

We found that the reaction lead to expected inflation was, in most cases, four quarters, but in other instances this turned out to be between two and four quarters. Batini and Haldane's (1999) dynamic simulations on a calibrated model of a small open economy show that the optimum lead length should be between three and six quarters. Amato and Laubach's (1999) similar attempt for the US economy find the optimum lead to be between five and eight quarters, but their basic interest rate rule involves no output stabilisation at the outset. Finally, Rudebusch and Svensson (1999) examine the performance of alternative policy rules on a small simulated model of the US economy, finding a broad support for forecast-based rules like [15].

In addition to that, we detected a substantial amount of interest policy inertia in all the countries we examine, as lags of the dependent variable are always found

[16] This point was kindly raised by R. MacDonald and S. Wren-Lewis.
[17] Results are not shown here for sake of simplicity, but are available from the author upon request.

significant. This too broadly agrees with Batini and Haldane's findings. This agrees with Woodford (1999), who argues that in presence of a zero lower bound on nominal interest rates and positive costs of inflation, a central bank's commitment to persistent interest-rate changes enhances social welfare.

Finally, as we have already mentioned, we explicitly tested for the possibility that the central bank might have responded to changes in intermediate objectives not included in our baseline specification [15]. The targeting of those additional variables is not the result of the fact that they provide information about future inflation and real activity. Such information is already contained in the estimated measures of expected inflation and potential output. In practice, changes in the policy instrument might instead be triggered by the desire to maintain fluctuations of the exchange rate vis-à-vis major trade partners within limited bounds, preventing excessive growth of domestic credit and liquidity aggregates, or shadowing the behaviour of some leading international interest rate. Clearly, as changes in such variables are often collinear with those in expected inflation, we would expect the addition of these regressors to lower the size of the estimated coefficient on expected inflation.

We now turn to discussing the existing empirical evidence on interest rate rules in OECD countries.

### *3.2 Monetary Policy Rules in OECD Countries*

The extensive literature on monetary policy rules has faced a relatively hard task when it came to envisaging testable models of its main hypotheses. This is probably because there are natural limits in relating the theoretical implications of contributions on central bank independence and credibility -largely of game-theoretic nature- to some empirical characterisation of policy behaviour. In fact, the first attempts to tackle this problem and translate some of the theoretical results of the political economy of macroeconomics (see Persson and Tabellini, 1999) into empirically testable propositions, produced relatively unappealing cross-sectional studies. Most often, these investigated the impact of institutions such as central bank independence and accountability on

[18] Batini and Haldane (1999) extensively elaborate on this point.

macroeconomic performance (Grilli, Masciandaro and Tabellini, 1991; Cukierman, 1992; Eijffinger and de Haan, 1996).

Subsequently, it became clear that the evaluation of observed monetary policy behaviour needed at the outset a clearer understanding of the various transmission channels of monetary impulses (Bernanke and Blinder, 1992). Following the identification of alternative measures to disentangle such channels, this empirical literature expanded along different lines.

First, many contributions have used Structural Vector Autoregressions to study the impact of alternative macroeconomic indicators on the policy stance, and the way in which policy impulses affect prices and the level of economic activity. This approach allows one to jointly model both the endogenous policy response and the effects it has on relevant macroeconomic variables, requiring relatively modest assumptions about the transmission mechanism. In this sense, Christiano *et al.* (1994), Bernanke and Mihov (1997, 1998), and Clarida and Gertler (1997) strongly contributed to refine such approach[19]. They in particular led to identify the monetary policy instruments in the US and Germany, over different samples, and using alternative orthogonalising structures for the fundamental macroeconomic shocks. Their methodology has become widespread, and its recent application to aggregate data for the euro area (Vlaar and Schuberth, 1999; Coenen and Vega, 1999; Tristani and Monticelli, 1999), testifies the flexibility of such approach in contexts of intense policy changes.

A subsequent strand in this broad approach used high-frequency, forward-looking data from financial markets to construct measures of unexpected shocks to monetary policy. Amongst these attempts, we signal Rudebusch (1995, 1996), and Bagliano and Favero (1999). The latter, in particular, derives exogenous measures of those shocks both in close- and open-economy contexts, and finds interesting evidence of simultaneity between German interest rates and the US dollar/German mark exchange rate.

A further recent study is by Favero and Rovelli (1999). In it, a model of the US economy , is estimated over the period 1960-1998. The output gap is first defined through a VAR specification, and inflation and a commodity price index are employed. In addition, the authors use GMM methods to estimate, over the period 1983-1998, an interest rate rule which allows to identify central bank's trade-off between output and inflation. Their approach requires full information, rational expectations, and invariance of

[19] See also Canova and de Nicolo' (1998).

the structural model to changes of the monetary policy regime. The results are in favour of the common belief (Clarida *et al.*, 1998) whereby since 1982 the Fed acted as a strict, though implicit, inflation targeter. This in turn rejects the hypothesis that the output gap is an independent argument in the policy reaction function[20].

There are two main problems with all VAR-based analyses of policy rules, though. First, the results seem to critically depend on the assumptions made about the transmission of shocks. Furthermore, once such identification restrictions are imposed, the estimation usually assumes a time-invariant structure for both the transmission process and the estimated policy reaction to economic shocks[21]. Second, as acknowledged, *inter alia*, by Christiano (1998) and Cochrane (1998), the interpretation of identified shocks and VAR estimated coefficients is particularly problematic in the case of policy rules. While VARs are ideal instruments to construct measures of monetary policy shocks for analyses of the transmission mechanism[22] (see also Gerlach and Smets, 1995), these models are much less useful when it comes to evaluating regime changes in the conduct of interest rate policy[23]. In periods of sustained financial innovation and structural change, the timing of the policy response and the transmission channels themselves are crucially affected. If the actual policy rule is of a forward-looking nature, as we proved is likely to be, the estimated coefficients of a VAR become of difficult interpretation. In addition, traditional VAR-estimated policy rules cannot easily become subject of structural change analysis, which is in turn crucial to investigate the eventual presence of policy shifts. The development of Bayesian and time-varying approaches to the estimation of vector autoregressions appears a much more promising avenue of research, and some research on their applications to the study of the transmission of monetary policy shocks is well under way.

Given these apparent caveats, it is no surprise that the latest strand of this literature has tried to deal with the issue from a single-equation, rather than from a system perspective (McNees, 1992, Groeneveld *et al.*, 1996; Muscatelli and Tirelli, 1996; Clarida and Gertler, 1997, Clarida *et al.*, 1998; Peersman and Smets, 1999; Gerlach and Smets, 1999; Gerlach and Schnabel, 1999).

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[20] As we shall see below, our results stand in sharp contrast with this.
[21] A similar severe drawback affects the interesting analysis in Broadbent and Barro (1997).
[22] For some cautionary views, however, see Rudebusch (1996), Bagliano and Favero (1998) Christiano et al. (1998).
[23] Clarida, Gali and Gertler (1999) wisely acknowledge how this point readily extends well beyond the limits of VAR-estimated policy functions.

In early studies, a key question was to devise ways of testing whether the adoption of inflation targets in countries like New Zealand, Canada, the UK and others, have had any significant effect on the overall credibility of the policy stance, or on economic performance. In general, results were not particularly supportive of a positive answer to such question, at least using data for the initial years of inflation targeting experiences (Freeman and Willis, 1995; Groeneveld *et al.*, 1996; Almeida and Goodhart, 1996). More recently, the focus of many studies has shifted towards testing the ability of simple interest rate rules to describe actual policy behaviour (Peersman and Smets, 1999). Gerlach and Scnhabel (1999), for example, find that a simple Taylor rule expressed in terms of aggregate average output gaps and inflation explains quite well the behaviour of average interest rates in EMU countries in 1990-98.

Clarida and Gertler (1997) estimate a reaction function for Bundesbank's short-term interest rate instrument, finding no apparent role for the growth of M3 in a simple rule expressed in terms of output and inflation objectives. Their approach is particularly interesting, because it represented the first explicit attempt allowing for a forward-looking behaviour on the part of the central bank, in so departing from conventional Taylor rules expressed in terms of lagged inflation[24]. Their overall conclusion is that Bundesbank's and Fed policies can be characterised as following a reasonably similar pattern.

Finally, Clarida, Gali and Gertler (1998) generalised the above approach, by estimating interest rate reaction functions for the G7 countries over 1980-1997, and innovating the measures of expected inflation and potential output used in Clarida and Gertler (1997). Their most general result, rapidly become a central tenet of the empirical literature, argues that the early 1980s marked a peculiar watershed in the conduct of policies in all those countries. Since then central banks appear to have turned invariably more aggressive in the use of interest rate changes in response to changes in expected inflation. More in detail, in a number of cases, and most crucially, in the US, monetary policy appears to have become gradually closer to a standard, though implicit, inflation targeting regime, in which output stabilisation concerns do not trigger systematic policy responses. In support to this argument, Favero and Rovelli (1999) claim that a significant output effect in an estimated reaction function might simply reveal that the central bank employs the current output gap as an indicator for future expected inflation. However, should this be the case, the output gap should be collinear with the adopted measure of expected inflation, or it should be able to anticipate inflation forecast errors. We were unable to find substantial collinearities between our measures of inflation and output gap (discussed below). In addition, the correlation with inflation forecast errors appears very limited and often has the wrong sign[25].

Quite clearly (see also Gerlach and Smets, 1999), in all policy analyses conducted within the framework of simple interest rate rules like our eq. [15][26], the results found critically depend on three factors. These are a) the methods used to identify measures of expected inflation and the output gap, b) the estimation techniques used to estimate the reaction function and to test them for structural breaks, c) the sample covered by such estimates. Since we believe that these issues radically affect the quality and interpretability of our results, we shall now devote some attention to illustrating and discussing the route followed in the present study.

### *3.3 Estimating Interest Rate Rules: Robustness Issues*

It is now customary to assume that, since the end of Volcker's 1979-1982 experiment, the growth of monetary base does not represent an adequate measure of the policy stance (Bernanke and Mihov, 1998; Christiano et al., 1994, 1998). This is true not just in the US, as Bernanke and Mihov (1997) and Clarida and Gertler (1997) have shown. Even when explicit money supply targets (reference values) are announced, as in Bundesbank's (ECB's) case (see chapter 3), central banks tend to react to changes in their final and intermediate objectives by eventually adjusting the price at which bank reserves are supplied to the interbank market. More in detail, while it is still controversial whether the policy stance can in all cases be characterised by the behaviour of a single interest rate instrument, the empirical literature almost universally agrees on the use of money market rates.

Throughout the last decade, monetary authorities have gradually increased the emphasis placed on repurchase operations, and correspondingly marginalised the traditional operations of discount window lending. Clearly, movements in money market rates are not exclusively triggered by policy actions, as demand conditions have some role

[24] Clarida and Gertler (1997) employed for expected inflation and potential output forecasts from a previously estimated VAR model of the German economy.
[25] These results are not shown here for reasons of brevity, but are available from the author upon request.

in determining short-run fluctuations in overnight rates like call money rates. Since mid-1980s, the operating procedures of the major central banks are such that shocks to demand for bank reserves are not entirely accomodated by monetary authorities, and thus trigger movements in the price of non-borrowed reserves. However, central banks's control over short-term money market rates critically depends on the signalling function exercised by various policy decisions, and these have usually greater impact at the shortest end of the term structure. Many recent attempts to empirically characterise reaction functions for monetary policy (Clarida, Gali and Gertler, 1998) use short-term money market rates as the policy instrument. We follow such route, and the exact choice of the interest rate for each country is briefly motivated in the Data Appendix. Our use of quarterly observations[27] should also represent a relatively safe way of filtering noisy short-run movements in call money rates, as at this frequency rates' movements should largely reflect authorities' policy stance.

Apart from the above dilemmas, the key assumption of any attempt to estimate simple monetary policy rules is the way in which the empirical model handles inflation expectations and central bank's information set. As we have seen, Favero and Rovelli (1999) results build on general assumptions of full information, rational expectations, and invariance of the structural model to changes in the monetary policy regime. In turn Clarida et al. (1998) employ a quadratic trend to obtain a measure of the output gap. As regards inflation expectations, the authors adopt an errors-in-variables approach to modelling rational expectations: future actual values are used as regressors instead of the expected values, and instrumental variable estimation is applied to account for the presence of forecast errors. Clearly, these choices, along with the fact that the reaction functions are estimated over a fixed sample, with no account for possible regime breaks and/or eventual structural changes, are stringent. In particular, they amount to assuming that:

a) Authorities, when producing forecasts of the level of economic activity, are fully informed about future output's DGP.
b) There is no scope for a learning process by policymakers about changes in the economic system, and by private agents about central bank's preferences. By consequence,

[26] Not necessarily, though, within the class of Taylor rules.

c) There is a time-invariant monetary policy regime throughout the sample.

Finally, our aim is to cover periods -the eighties and nineties- during which both institutional *and* structural change, despite the assumptions held by previous analyses of monetary policy, was sustained. In particular, following Stock and Watson (1999), we suspect that the inflation and output processes in all the countries we study have undergone significant structural breaks. We thus decided to explicitly test for such belief: its empirical validation would yet again confirm the need for a limited information approach to our problem.

There is now a wide range of technical contributions devoted to study trend breaks in unit roots and the problems associated with the endogeneity of break points[28]. Here we limit ourselves to study whether output and inflation DGPs have undergone major shifts over our sample period, and leave aside the determination of the exact number and position of the break points. We employ the class of tests devised in Hansen (1992b), and Andrews and Ploberger (1994). They all belong to the broad category of Chow-type tests with unknown break point(s), and build upon the assessment of the significance of the value of the LR, Wald and LM statistics derived from recursive switching regressions. We use the *MeanF*, *SupF* and the $L_c$ variants advocated by Hansen (1992b). The first two tests have parameter constancy as their null against the alternative of sudden breaks, whereas the latter statistic is for the alternative of a smooth change. In our case, the testing strategy requires prior estimation of univariate models for output and inflation. We adopt simple autoregressive specifications including trends and constants when required, and the semiparametric, fully modified FM estimator of Phillips and Hansen (1990) and Hansen (1992a). The latter is a two-step methodology that first estimates the asymptotic covariance matrix of the system, and then provides regression parameters[29]. Tests on the null of parameter stability are finally carried out[30].

The test statistics are reported in **Table 1** for all the countries we examine. In the case of inflation, the sample covers the years 1971Q3-1997Q4, whereas for GDP some data constraints for Sweden and New Zealand substantially shortened the sample we used. The first column of each section displays the estimated statistic for the three tests. As

[27] Clarida *et al.* (1998), instead, employ monthly data.
[28] For an extensive account of this debate, see Stock (1994).
[29] Additional details on the testing procedure, as well as on the results we summarise here, can be obtained from the author upon request.

regards output, stability is rejected at conventional significance levels in all countries but Japan. In the case of inflation, the results are even more clear-cut, with all countries displaying instability at very low significance levels. These estimates are robust to changes in the kernel and bandwidth parameter chosen to filter the residuals, as well as to alternative functional forms for the specified models.

The above results strongly support the need for a modelling approach that allows unobservables to be estimated according to some time-varying pattern. We thus decided to explicitly rely on the assumption that the private sector is imperfectly informed about the central bank preferences, and that the central bank is imperfectly informed about the permanent and cyclical components of output growth (see Orphanides, 2000). Both points seem more in line with a forward-looking view of the interactions between policymakers and private agents, while taking into account the limited information available to central banks about the working of the economic system (Blinder, 1998).

The Kalman filter and Structural Time Series techniques (STS, Harvey, 1989) we employ for generating the necessary measures of expected inflation and potential output, represent a natural tool to take into account the limited information and time-varying nature of the policy process.

As regards potential output and aside from the above considerations, standard filtering techniques like Hodrick-Prescott or Baxter-King are usually associated with several drawbacks. In the former especially, the choice of the trend-smoothing parameter depends on the variance ratio of the shocks to all stochastic components, and it is then highly sample-dependent, other than relatively arbitrary. An interesting approach is undertaken in Gerlach and Smets (1999), who study the behaviour of output gaps in the EMU area using an unobservable-component method. Their model is however pretty complex, and would certainly bear substantial costs in terms of the subsequent estimation of reaction functions. We do however believe that Gerlach and Smets' (1999) techniques are broadly in line with the STS approach that we use (Hamilton, 1994; Maddala and Kim, 1998; Kim and Nelson, 1999).

There are various advantages in using such methods. First, they provide a simple way of decomposing the series we observe into trend and cyclical stochastic components, which is particularly convenient when estimating unobservables like potential output and expected inflation. Second, STS models are parsimonious models that however have

[20] Estimates were conducted by adapting a GAUSS code kindly provided by Bruce Hansen. Hansen (1992b) tabulates

peculiarly rich ARIMA processes as their reduced forms (Harvey, 1989). Finally, this modelling approach (see also chapter 4) is implemented by applying a Kalman filter algorithm, which is a pretty natural way of accounting for a gradual learning process by policymakers and private agents.

The way in which we obtained measures of the two unobservables for each country is as follows. We estimated quarterly models for real GDP and inflation for each country, obtaining a decomposition of the series into trend, cycle, and irregular components.[31] In the case of GDP, a convenient decomposition of the series was generated by applying the Kalman filter on the trend component. The latter was then computed based on one-step-ahead predictions of the state vector. This way, estimates of potential output are based only on past information, rather than on the full sample.

In the case of inflation, we simply computed one-step-ahead prediction errors from a univariate STS model to obtain a measure of expected and unanticipated inflation. This way, model's parameters are updated only gradually, as new information become available. The use of the basic filter, as opposed to the smoothing algorithm (Kim and Nelson, 1999), guarantees that future observations never affect the calculation of the stochastic components.

Formally, the general class models we estimate is the following:

$$Z_t = \tau_t + \omega_t + \varepsilon_t \qquad [16]$$

where $Z_t$ is either inflation or output, $\tau$ and $\omega$ are the trend and cyclical components, and $\varepsilon$ is a random shock. In turn, the trend component is specified as

$$\begin{aligned} \tau_t &= \tau_{t-1} + s_{t-1} + \vartheta_t \qquad & \vartheta_t = NID\left(0,\sigma_\vartheta^2\right) \\ s_t &= s_{t-1} + \zeta_t & \zeta_t = NID\left(0,\sigma_\zeta^2\right) \end{aligned} \qquad [17]$$

where $\tau$ represents the actual value of the trend and $s$ its gradient.

In addition, both real GDP and inflation contained marked cyclical, non-seasonal components. We modelled these by estimating the series with one or two stochastic

asymptotic critical values for each of the tests performed here.

[31] The STAMP 5.0 software was used to estimate the STS models, through the conventional concentrated diffuse likelihood technique. Output and inflation were found to be I(1), and to have significant cyclical components. For a similar approach to forecasting inflation in the presence of potential structural breaks, see Stock and Watson (1999).

cycles, as appropriate. These stochastic cycles are defined recursively as follows (Harvey, 1989):

$$\begin{bmatrix} \omega_t \\ \omega_t^* \end{bmatrix} = \rho \begin{bmatrix} \cos\lambda_c & \sin\lambda_c \\ -\sin\lambda_c & \cos\lambda_c \end{bmatrix} \begin{bmatrix} \omega_{t-1} \\ \omega_{t-1}^* \end{bmatrix} + \begin{bmatrix} k_t \\ k_t^* \end{bmatrix}, \qquad [18]$$

where $\lambda_c$, $0 < \lambda_c < \pi$ is the frequency, in radians, the $k$s are white noise uncorrelated shocks, and $\rho$ is a damping factor.

In our case, we assumed the slope of the trends as following a stationary, first-order autoregressive process[32].

To see whether our models could be improved by extending the information set, we tried also multivariate STS specifications for the two processes. In the case of GDP, additional regressors did not provide a better fit than our univariate specifications. In the case of inflation, we examined whether lagged values of variables such as the exchange rate, output growth, short-run interest rates and the money supply could help to forecast future inflation. In all cases, the benefits of extending the models seemed to be quite modest. In part, this is due to the fact that the univariate representations are more parsimonious, while a detailed *ad hoc* specification search would have not led to dramatically different measures of expected inflation.

**Figure 1** compares our measure of the output gap with that obtained from a H-P filtering procedure for the USA[33]. It shows that our measure differs markedly from that used in previous studies, and indeed that quadratic or H-P trending procedures tend to exaggerate the cyclical component.

**Figures 2** and **3** plot our measures of (4-quarter ahead) expected inflation and implied ex ante real interest rates for the G-3 countries and the four inflation targeting economies, respectively.

We now finally turn to the illustration of our empirical findings. The reaction function in equation [15] is estimated for each country using simple Recursive Least Squares[34]. Recursive estimates are particularly useful when, as in the present case, one

[32] I.e., the trends were specified according to a "stochastic level, damped slope" formulation, which however did not yield substantial differences relative to the "fixed level, stochastic slope" case.
[33] Plots of the output gaps and expected inflation for all other countries in our sample are not presented here for space reasons, but are available upon request. Our estimates correspond well with descriptive accounts of macroeconomic conditions in the countries under consideration. Fitting a quadratic trend, as in Clarida *et al.* (1998) produces a even more marked cyclical pattern than the H-P measure depicted in Figure 1.
[34] Estimated coefficients are obtained computed with a GAUSS code. Stability tests are conducted using PcGive 9.1.

needs to observe how estimated coefficients evolve over time. Moreover, we perform conventional structural-stability tests on the residuals of each equation, to capture possible signs of breaks.

We estimate the quarterly models for the G-3 economies over samples starting in 1970 and ending in 1997. Next, sub-sample estimates are computed, to assess whether and how results are affected by the dominance of particular events or regimes over specific periods. We do the same for the four inflation-targeting economies we chose, but in their case, the available output and interest rate data do not go further back than early 1980s.

The data we use are quarterly observations taken from OECD Main Economic Indicators and IMF International Financial Statistics, as described in the Data Appendix.

The next two sections describe our major results for each of the two groups of countries, starting first with the G-3 economies.

## 4. Monetary Policies in USA, Japan and Germany: Results from Estimated Interest Rate Reaction Functions

According to narrative accounts, monetary institutions in the G-3 (the U.S., Germany and Japan) have been remarkably stable during the sample period. In particular, the relationships between political systems and monetary institutions have not undergone significant changes, if one excludes from this definition the German re-unification[35]. As is well known, in the U.S. and Germany the central bank enjoys a relatively high degree of independence (see Cukierman 1992; Eijffinger and de Haan, 1996) and is best defined as a "goal independent" central bank[36], that is, a bank which is not held accountable for achieving a certain policy target.

**Figure 2** shows that 1979 was clearly a turning point for US monetary policy, as real rates tend to become positive only after that date. Closer to our days, monetary policy during the Greenspan era has been defined as "pre-emptive monetary policy without an explicit nominal anchor" (Mishkin, 1997).

In the views of some scholars, German monetary policy over the eighties and nineties appears as a regime of "disciplined discretion" (Laubach and Posen, 1997). In

[35] Since 1979, EMS membership might have constrained the Bundesbank's ability to retain control of monetary policy. Most discussions on the DM's role in the EMS have concluded that the Bundesbank largely retained its independence (see Fratianni and von Hagen, 1990; von Hagen, 1999).

[36] See Fischer (1995) For instance, Neumann (1996) and Clarida and Gertler (1997) argue that, while the Bundesbank was pursuing multiple objectives, it retained considerable flexibility as to how to achieve them, in the sense that emphasis sometimes shifted from one policy target to another. For a similar view see Mishkin and Posen (1997). For a contrasting view, stressing continuity in the Bundesbank's use of monetary targets, see Issing (1997).

fact, the mid-eighties witnessed a period of generally restrictive policies, and this was partly due to the great variability of the exchange rate with the dollar, and to shifts in the terms of trade. German re-unification also created a major challenge, as the Bundesbank engineered a rapid increase of real interest rates in an effort to control inflation.

The overall picture for Japan is somehow less clear-cut, as financial instability and sharp exchange rate fluctuations influenced interest rate management in opposite directions. Despite this, the picture we obtain from the estimation of its reaction function does yield results in line with some previous findings (Chinn and Dooley, 1997).

Results for these three countries are reported in **Tables 2-4**. For ease of exposition, we list for each country only the long-run static solutions of the model, as each regression contains one or two lags of the dependent variable. Asymptotic standard errors are reported below each estimated coefficient, and summary statistics of the regressions are included.

When we estimate the **USA** reaction function over the whole sample period (**Table 2**), we find that the coefficient on expected inflation is not significantly larger than one, and detect a coefficient associated to the output gap overall not very significant. Diagnostics tests and recursive graphs[37] show marked instability before 1985, culminating in 1979-1982's experiment of monetary base targeting, which involved greater instability in money market rates. Since then, the Fed has opted for the targeting of money market (federal funds) rates. Goodfriend (1995), Bernanke *et al.* (1999), Clarida *et al.* (1998), and many others argue that this parenthesis also marked Fed's change of attitude towards the appropriate degree of aggression on inflation expectations.

When re-estimated over the post-1980 sample, the US reaction function confirms that some important changes did indeed take place. Plots of the recursively estimated (long-run) coefficients and error bands are shown in **Figure 4**[38], along with 1-step up and N-step down Chow tests for structural stability. Interest rates now seem to react to inflation expectations on a shorter horizon (a 2-quarter horizon is found to work best post-1985) and with a larger coefficient than over the whole sample. These results are at odds with those obtained by Clarida *et al.* (1998)[39], as they detect an estimated coefficient on inflation that is much greater than one. The most likely explanation for this difference

[37] Here shown only for a shorter sample.
[38] These were computed using the author's GAUSS routine and plotted using GiveWin 9.1.
[39] Mehra (1997) estimates a somewhat atheoretical reaction function, where the Fed funds rate follows an error correction process and responds to the output cycle and to the interest rate on long term treasury bills. We added the latter variable to our equation, but could not find any significant effect.

seems to lie in their fixed sample period, as we found that the size of our estimated inflation coefficient depends critically on the sample we choose. In fact, the picture changes again when we focus on the post-1985 sample. The equation is very stable, and includes a coefficient on expected inflation with a point estimate greater than unity (although it is not significantly larger than 1). The post-1985 reaction function seems to suggest that the Fed was adjusting real rates to follow the output cycle, with Figure 4 showing a significant output gap effect by 1992. One might argue that having successfully clamped on inflation expectations since mid-eighties, the Fed exploited its reputation to implement countercyclical policies. Furthermore, the theoretical model discussed above suggests that in a full information context, that is, when the private sector has learned about the bank preferences, inflation expectations might be collinear with the output cycle. This might bias the estimated coefficient on inflation expectations downwards. The remaining interesting aspect of the post-1985 results is that they show a shorter lead on expected inflation (2 quarters) than in most of our other estimated reaction functions.

Our estimates for **Japan**'s reaction function **-Table 3-** over the whole sample show a not significant coefficient on the output gap, while that on expected inflation is significant but well below one. Furthermore, the equation performs poorly, as illustrated by the diagnostic tests. We tried to improve on this by including some additional regressors. As it turns out, the US Federal Funds rate exerts a strong influence on Japanese policy. As in the case of the US, however, instability in the reaction functions persists in the 1980-82 period. Shortening the sample to the post-1982 period results in a dramatic increase in the expected inflation coefficient, which suggests that central bank's attitude towards inflation changed markedly. On the other hand, the recursive estimates -**Figure 5**- show that cyclical conditions became important only after 1992. The structural stability tests show that there is likely to be a large break around 1986. This was probably due to external pressures exerted on Japanese monetary policy in relation to the G-7 agreements on the value of the US$. It also confirms the casual observation that Japanese monetary policy might not have been sufficiently geared towards domestic targets (see *The Economist*, July 17, 1998), and that this might have contributed to excessive deflation in early 1990s.

**Table 4** reports the estimated reaction function for **Germany**, for the full sample period and since 1980. The estimates for the whole sample show that interest rates reacted to inflation expectations (with a point estimate greater than 1) and output. The addition of the US Federal Funds rates marginally improves the fit of the interest rate reaction function[40].

The variable addition tests show that neither money growth nor the exchange rate (measured as the DM-US$ rate) seem to exert an independent significant effect on German interest rates. This is interesting and confirms the results in Clarida and Gertler (1997), and Bernanke and Mihov (1997). Since 1974, the Bundesbank set target ranges for the growth of broad monetary aggregates, but over the fifteen years preceding the start of Stage Three of EMU, actual growth rates often exceeded (fell short of) the upper (lower) limit of the targeted band[41]. This confirms most accounts of Bundesbank's policy stance. Monetary targets were not the Bank's primary objective, and discretionary undershoots and overshoots of the target bands were allowed where this did not impair the achievement of the inflationary objective.

The diagnostic tests for the estimated model show some signs of non-normality (and possibly ARCH) in the residuals, but this is due to the bunching of a small number of large residuals at the end of the 1970s.

The estimated reaction function for Germany is overall remarkably stable, with the estimated coefficients constant across sub-samples. **Figure 6** shows 1-step up and N-step down Chow tests, as well as the estimated coefficient and standard error bands of the expected inflation and output gap regressors for the post-1980 regression[42]. This confirms the stability of Bundesbank's policy rule. It also shows that the size of the estimated response to output gap fell after the unification shock in 1991, as the Bundesbank tried to offset the subsequent inflationary shock. We still find that a four-quarter lead for expected inflation works best for the post-1980 sample.

To sum up, two main points emerge from our estimates of the German function. First, the relatively good performance of the estimated interest rate reaction function suggests that the underlying policy objectives were remarkably stable across the whole sample. Second, there is some evidence supporting the view that the policy thrust turned

[40] For a descriptive account of these effects see Mishkin and Posen (1997).
[41] See von Hagen (1993, 1999), Issing (1997).
[42] It is worth noting that the estimated coefficients only show the short-run effect and do not take into account the impact of the autoregressive component of the reaction function. However, the estimated coefficient on the autoregressive term

gradually more conservative since mid-eighties and even more after the re-unification. Moreover, in line with recent work (Clarida *et al.*, 1998) we find that monetary policy in Germany reacts systematically to cyclical conditions, even though the Bundesbank's declared monetary strategy (see Issing, 1997) was entirely expressed in terms of monetary targets.

The analysis of Germany's monetary policy also highlights some noticeable differences with Fed's policy behaviour. Although both estimated reaction functions seem substantially stable post-1985, the Bundesbank appears to respond more aggressively to movements in expected inflation than the Fed. Alternative interpretations are, however at hand. Mishkin and Posen (1997) label Fed's policy as "just do it", or "pre-emptive policy without a nominal anchor". Their argument is a classical one whereby policies must act well in advance of a surge in inflation expectations, as the full impact of monetary policy on inflation takes long lags. The main drawbacks of such policy obviously lie in the difficulty of establishing a clear policy pattern, with all the risks that this implies at times when the economy is hit by major exogenous shocks. Our results partly support such pragmatic and forward-looking attitude for the Fed. This does not translate, however, as if the Fed systematically reacted to longer-term expectations, as in Bundesbank's case. In fact, we found that shorter leads on the expected inflation regressor (two instead of four quarters) seemed to work better in the case of the US over the latter part of the sample. This confirms the casual observation that the Fed has chosen to signal its commitment to low inflation only in recent years, by reacting in advance to increases in inflationary expectations. Overall, Clarida *et al.*'s (1998) finding, of a US reaction function behaving *as if* the Fed operated according to an implicit inflation targeting framework, are only partly supported by our findings, and only since early 1990s.

The general conclusion is that the G-3 policy reaction functions look relatively different from one another. In addition, despite having overall stable monetary institutions, policy rules in the G-3 seem to have smoothly evolved along alternative lines. German interest rate policy appears to have become more conservative after the re-unification. In Japan, it seems to have been strongly influenced by external objectives, until very recently. In the US, the highly successful countercyclical monetary policy of the Fed seems to be purely a very recent *datum*. As the existing empirical literature employed full-sample, full-information estimation techniques, our findings appear peculiarly valuable.

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is quite stable, and the recursive coefficient estimates also reflect the movement in the long-run coefficient for each

## 5. Policy Rules under Inflation Targeting: United Kingdom, Canada, Sweden and New Zealand

For most of the sample period, central banks in this second group of countries enjoyed limited independence in the conduct of their policies, at least in comparison to those of G-3 countries (see Cukierman, 1992; Eijffinger and de Haan, 1996). During the 1990s explicit inflation targets were announced in all countries, but there are relevant differences between the institutional arrangements set in each experience. For instance, only in New Zealand (and to a slightly different extent in the UK since May 1997), the central bank has a legal mandate to achieve the target.

**Figure 3** plots the expected inflation series and the *ex ante* real interest rates computed using our expected inflation series for the group of inflation targeting economies. Interestingly, in the case of Sweden, Canada and New Zealand, *ex ante* real rates appear to have turned positive well before the announcement (represented in the charts as a vertical solid line) or the adoption of targets. In addition, inflation expectations, at least in the case of the UK, Sweden and New Zealand, seem to have been somewhat subdued before the announced regime changes. The regime change seems to have simply consolidated some prior gains in terms of lower inflation.

In the **United Kingdom**, the Bank of England was granted operational independence only in May 1997. However, several changes affected UK's monetary strategies in recent times. The election of the Thatcher government in 1979 signalled a long-lasting shift in the collective attitude towards inflation[43]. Instead of adopting an institutional approach, the conservative governments tried to build a reputation for their commitment to low inflation policies, envisaging with the MTFS, amongst other things, a 5-year sequence of gradually decelerating growth targets for £M3. However, the unstable relationship between this monetary aggregate and the final policy objectives quickly led to the demise of formal monetary targets. The government then adopted a more eclectic approach (see Minford, 1993, King, 1998), which essentially involved targeting nominal income growth. In the late 1980s, the exchange rate assumed greater importance as an indicator of monetary conditions (see Bowen, 1995), and Sterling finally entered the ERM

explanatory variable.

[43] Alogoskoufis et al. (1992) find convincing evidence of a spectacular reverse in the political business cycle after Mrs. Thatcher came to power. For a more descriptive analysis, see Minford (1993) and Bowen (1995).

of the European Monetary System in 1990. The exit from ERM following the 1992 crisis forced the government to put an alternative regime in place, and the post-1992 announcement of explicit inflation targets was seen as a practical way of achieving price stability. However, the central bank played only the role of publicly assessing the overall consistency of the policy stance. The newly-elected Labour government in 1997 then sought to further enhance the inflation targeting framework, by granting the Bank with instrument independence (for a definition, see Fischer, 1995). Monetary policy decisions are now taken by a recently-constituted Monetary Policy Committee.

Our estimates for the UK -**Table 5**- show that over the whole sample period the coefficient on inflation expectations is not significantly larger than one. Furthermore, the money market interest rate seems to have reacted to both the exchange rate and money supply.

Given the instability in the estimated reaction function until the mid-1980s, we re-estimated the equation for the 1980-1996 sample, and then we further shortened the period to 1985-1996[44]. Results show that the policy horizon became substantially shorter after the 1985 sterling crisis, as interest rates appear to react to one-quarter ahead expected inflation, and the coefficient on expected inflation becomes significantly larger than one. Along with this, other minor shifts in policy regimes are also apparent (**Figure 7**). For instance, the estimated coefficient on the sterling effective exchange rate was significant between 1988-1992, capturing both the 'shadowing the DM' and the ERM phases in UK policy (see also the behaviour of the *t*-ratio for the estimated coefficient). By contrast, the coefficient on the output gap became permanently less significant during the ERM phase, as domestic output objectives were sacrificed for the external objective.

These findings closely mirror the changes in policy regimes outlined above. The main turning point is in 1979. The more recent shifts in the estimated coefficients of the reaction function seem to be linked to the difficulties encountered in achieving a specific target rather than a lack of commitment to the goal of price stability.

Since the breakdown of M1 as an intermediate target in early 1980s, until 1991 the Bank of **Canada** had not committed itself to any pre-determined policy pattern, aside from the reiteration of a long-term goal of price stability. Neither intermediate target, not

[44] In this and all other cases, extending the sample to include 1998 and 1999 could have helped understanding recent developments. However, the great interest rate instability associated with the Asian crisis could have heavily affected the full-sample results. We however defer the assessment of such events to some future occasion.

time frame was apparently cast in the attempt to pursue the long-run objective, while various monetary and credit aggregates (including the exchange rate with the US$) were used in turn as information variables. In 1991, the government and the Bank set a sequence of year-to-year target bands for the inflation rate, to bring in a gradual reduction in inflation. However, the central bank was not granted a legislative mandate to achieve those targets, nor was a procedure established whereby the bank would be held accountable for missing the targets. The "doctrine of dual responsibility" traditionally attributes the ultimate responsibility for the results of monetary policy to the Minister of Finance. Thus, the Bank of Canada has enjoyed only a limited degree of formal independence (see Cukierman, 1992; Eijffinger and de Haan, 1996). Nonetheless, the monetary authorities had been publicly calling for a stricter overall control on inflation as early as 1988, while since 1994 the degree of transparency and accountability of their acts has remarkably increased (Mishkin and Posen, 1997).

Our estimates for Canada over the full sample period (1975-1997) yield somewhat puzzling results (see **Table 6**). When the US Fed funds rate is added to the equation, both the coefficients on the output gap and on expected inflation become irrelevant. Clearly, as in the case of Germany and Japan, the Fed funds rate absorbs part of the significance of the inflation regressor. Although M1 growth was the intermediate policy target in Canada between 1975 and 1982[45] (Freedman, 1995), we could not find a significant role for money supply in our estimated reaction function. Furthermore, there are clear signs of instability in the estimated function in the late 1970s and early 1980s.

Once the equation is re-estimated over the post-1982 sample, we find that the coefficient on inflation expectations is still not significant, whereas effective exchange rate changes now seem to be relevant, along with the Fed funds rate.

What about the impact of inflation targets? The announcement of targets, which took place in early 1991, does not seem to coincide with a break in the behaviour of interest rate policy (see **Figure 8**). At most there seems to have been a temporary impact on interest rate policy *just prior* to the introduction of inflation targets, as some signs of instability in the expected inflation coefficient are detected around the period 1990-1[46]. Descriptive accounts of Canadian monetary policy in this period (Mishkin and Posen, 1997) point out that the inflation target was mainly used as guidance for expectations. They also stress that in several occasions monetary policy was in fact constrained to react

[45] In 1982, it was officially abandoned due to innovations in the financial sector.

to external conditions, such as exchange rate developments and the behaviour of US monetary policy. Our estimated reaction function seems to confirm this. Furthermore, the Bank has recently defined a short-run operational target, the index of monetary conditions (MCI). MCI changes include variations in a short-term interest rate and in the trade-weighted exchange rate. Clearly, this highlights the importance of external constraints on the Bank of Canada's policy stance.

Since 1977, **Sweden** had been unilaterally pegging its currency, first to a trade-weighted basket of currencies, then switching to the ECU in May 1991. However, the attitude whereby this commitment to the external anchor was pursued varied significantly, as numerous devaluations took place (Horngren and Lindberg, 1994). To some extent, the Riksbank turned to a less accomodative stance towards inflation developments after 1982. The marginal (overnight) rate was then extensively used to regulate large currency flows during the fixed-exchange rate period. After the 1992 crisis, the Riksbank floated the krona, and then announced the unilateral adoption of an inflation target in January 1993[47]. However, the bank has never been granted an independent status, and political influences on the board appear important (Svensson, 1995).

The full-sample estimates (1982-97) for Sweden show a significant but relatively low coefficient on expected inflation, while the output gap is not significant at all (see **Table 7**). The main instability in the estimated reaction function corresponds to the time of the ERM crisis in 1992, when the krona was forced to devaluate with respect to the ERM parity despite an unprecedented surge in domestic interest rates. Since then, Sweden has adopted inflation targeting. However, Svensson (1995) points out that the credibility of the new regime has been hampered by a number of factors, such as the deep political divisions over the conduct of monetary policy and the relatively large budget deficits. The sudden policy reversals and the overall uncertainty about the post-1992 regime clearly show up in our estimates, making it difficult to detect a clear policy pattern.

Once a dummy for the ERM crisis in 1992 is included, the coefficient on expected inflation rises and becomes more significant, but the point estimate remains below one. The output gap variable is almost significant at the 5% level. However, we are unable to find signs of a significant permanent shift in the reaction function following the

[46] Although the N-step down Chow tests are not significant at the 5% level

[47] The term unilateral emphasises the lack of a legislative mandate to achieve a specific inflation target. See Svensson (1995) for a detailed account of these events.

introduction of inflation targets. The main fact that emerges from **Figure 9** is (as for the UK) the decreasing importance of domestic inflation and output targets just before the ERM crisis in 1992. On the other hand, since inflation did in fact fall in Sweden, one might conclude that monetary policy in this period mainly acted to keep real interest rates high until inflation was brought down. Taking into account the severe credibility constraints outlined above, this apparently stubborn policy was perhaps the only alternative left to the bank in order to signal its willingness to curb inflation.

Finally, we turn to the evolution of the monetary regime in **New Zealand**, which switched to inflation targeting in 1989. Historically, New Zealand's Reserve Bank had a degree of independence that ranked lowest amongst the OECD countries (see Cukierman, 1992; Eijffinger and de Haan, 1996). Correspondingly, New Zealand's inflation rate was well above the OECD average. Up until the mid-1980s monetary policy relied mainly on regulation and administrative controls of capital markets. Since 1985, the Bank has turned to a more market-oriented approach to monetary control, and based policy decisions on a variety of indicators, such as the exchange rate, the term structure of interest rates, monetary aggregates and output (see Fischer and Orr, 1994). The Reserve Bank Act, introduced in 1990 to establish a legislative commitment to price stability, gave the Government and the Central Bank Governor the mandate to agree on a policy target (it was decided that this should be an inflation target). The Act explicitly contemplates the possibility of the Governor's dismissal if the set target is not met.

For these reasons, New Zealand has been the most-often-quoted inflation-targeting experience. This not least because the legal arrangements designed to regulate bank's activity follow the prescriptions of policy design theory more closely than elsewhere (see Walsh, 1995, 1998). The estimated equation for the full sample (see **Table 8** and **Figure 10**) shows that interest rates seem to have reacted only to expected inflation - the estimated coefficient is close to be significantly larger than 1 - whereas domestic cyclical conditions do not seem to matter much[48]. Although exchange rate shocks are explicitly quoted in the Bank charter as a possible justification for deviating from the announced policy, we could not find a significant exchange rate effect. On the other hand, diagnostic tests signal some ARCH pattern in the residuals. This may be due to occasional

[48] Hutchison and Walsh (1998) suggested that the Reserve Bank looked at output stabilisation as an additional objective, but the output gap term is not significant in our estimates. Nevertheless, as pointed out previously, the absence of an output gap term in the reaction function does not preclude some degree of output stabilisation.

interest rate adjustments to external conditions. Another possible explanation can be found in the relatively narrow band originally set around the inflation target, which caused significant instrument instability in a futile effort to "fine tune" inflation control[49] (Mishkin and Posen, 1997). Again, the essential result from the stability tests is that in the '90s the Central Bank followed a policy pattern that had already been established in the former decade. It is impossible to detect significant breaks in correspondence of the announcement of inflation targets. The other main point to note is that inflation targeting does not seem to have provided the authority with a greater leeway to stabilise output fluctuations. The stability of the coefficient attached to inflation expectations and of the overall equation indicates that the inflation target regime does not seem to have made a marked difference to interest rate policy.

## 6. Concluding Remarks

In this chapter, we have discussed and estimated forward-looking interest rate reaction functions for two groups of OECD economies. Our aim was twofold. First, we sought to envisage whether the recent emphasis placed by the existing empirical literature on the consistency of monetary regimes in the G-3 economies with an inflation-forecasting approach, was justified. Second, we wished to establish empirically whether there was any systematic pattern between institutional change, in the form of the adoption of explicit inflation targets and central bank reforms, and the operational conduct of monetary strategies. In addition to the detailed results for each country set out above, a number of general conclusions emerge from our empirical results.

First, with the exception of Germany and the UK (since 1992), most of the monetary authorities in our sample do not seem to follow stable simple forward-looking policy reaction functions based on output gaps and expected inflation (and, *a fortiori*, Taylor rules). This suggests that caution has to be exercised in using an inflation-forecast targeting framework to model monetary authorities' preferences (see Clarida *et al.*, 1998; Favero and Rovelli, 1999). In the US and Japan, countries where there have been no major central bank or other institutional reforms, we find that policies did evolve to a considerable degree in the 1980s and 1990s. However, it is only since the 1990s that

[49] Perhaps not surprisingly, both the inflation target and the band width were revised in the '90s

estimated interest rate rules in these countries begin to look like the ones the theoretical research on the inflation-forecast targeting approach has recently illustrated.

Second, in countries where there were explicit intermediate targets (such as the growth of M3 in Germany), these appear mainly (see also the discussion in Chapter 3) as a device to anchor expectations. In practice, policy is not constrained to follow them strictly. In addition, monetary policy is often found to follow a broader set of objectives. Our results confirm those of previous researchers who have detected in the Bundesbank a marked "targeting" attitude regarding inflation, output, and some external conditions. More generally, where the policymaker is subject to implicit external constraints (as in the case of Canada, Japan, and to a lesser extent, Sweden), this can sometimes lead to a less interpretable picture. This sends a negative signal regarding the ability of simple interest rate rules expressed in terms of inflation and cyclical conditions in capturing monetary policy changes.

Third, with the exception of the UK, the recent switch to inflation targets in the countries we studied does not seem to have radically altered the way in which interest rate policy reacts to changes in its final objectives. In practice, there is some evidence, particularly clear in Canada's and New Zealand's cases, that any major changes in the responsiveness of interest rates to expected inflation took place well before the adoption of inflation targets. The same pattern seems to have been followed even when such institutional reforms have been accompanied by greater central bank independence. A possible interpretation is that the new regimes were brought in simply to consolidate gains in terms of lower inflation. Only longer datasets will tell whether, in response to major exogenous shocks, monetary policy will be able to respond more vigorously to inflationary forces than in the past.

Finally, we detected some important differences in the behaviour of central banks as far as output stabilisation is concerned. On the one hand, we found some evidence in favour of an apparent 'just do it' attitude of the Fed. That is, the central bank, at least since 1990 seems to exploit its consolidated reputation to focus on the cycle. This pattern is reflected, to some extent, by the shorter optimal lead placed on expected inflation in the estimated reaction function. At the other extreme, some monetary authorities apparently feel the need to build up a reputation. This was particularly clear in the Swedish Riksbank's stubborn attempt to lower inflation expectations by means of high interest rates and the apparently exclusive focus of the Bank of New Zealand on domestic inflation.

Whether this 'reputation-building' phase will also apply to those central banks that have only recently acquired their independence, such as the Bank of England and the European Central Bank, remains, however, a fairly open question.

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## Data Appendix

The data we used were quarterly series, extracted from OECD *Main Economic Indicators*, apart from a few cases, in which the source is equivalently quoted. In most cases, we were able to employ seasonally adjusted data.

For each country, we measured output using the GDP at constant price series. For Sweden and New Zealand the available constant price series for GDP do not go back further than 1980 and 1982Q2, respectively. The inflation series were defined as simple 4-quarter log-differences in the all-items CPI, except for Britain, where it was the equivalent change in the index of retail prices excluding mortgage interest payments (not available before 1975).

The index of effective exchange rates (trade weighted) was the measure for the exchange rates. Also, spot exchange rates vis-a-vis the US dollar were tried for Japan, Germany, Canada, New Zealand and the UK; vis-a-vis the German mark for the UK and Sweden.

The rate on US Federal Funds was used as the foreign interest rate for Japan, Germany, Canada, and New Zealand. The 3-month FIBOR German rate was the foreign rate for the UK and Sweden.

Below we briefly outline the short-term interest rates we chose as policy indicators, along with the monetary aggregates we applied in the generation of regressors. The rates are generally converted from monthly series.

| *Country* | *Modelled Interest Rate Variable* | *Money* |
|---|---|---|
| *USA* | Federal Funds Rate. As noted in the main text, during the early to mid-80s, the FFR provides an accurate measure of Fed's policy stance. The only exception is the Volcker experiment in the 1979-82 period, when operating procedures could be better summarised by a different instrument choice (inter alia, Bernanke and Mihov, 1998; Goodfriend, 1995) | M1 |
| *JAPAN* | The Call Money Rate (rate between financial institutions) is directly affected by Bank of Japan's reserve management policy, through discount window and open market operations (see Ichimura, 1993) | M2 plus CD |
| *GERMANY* | Bundesbank's intentions are mainly reflected by the rate in the market for interbank reserves, the Call Money Rate. In facts, the discount window lending to commercial banks exclusively affected the behaviour of this rate until 1985, when the banks started to be supplied with reserves through repurchase operations. Since then the call rate shadows the rate on these loans (REPO rate). (see Bernanke and Mihov, 1998; Clarida and Gertler, 1997) | M3* |
| *UK* | We use an Overnight Interbank Rate series post-1983. This is not available pre-1983, and thus we employ the Rate on 90-day Treasury Bills, which displays a very close correlation with the interbank lending rate, for those observations (source: IMF, IFS). | M4 |
| *CANADA* | The Bank of Canada introduced in 1996 the concept of Monetary Conditions Index (MCI) as its short-run operational target. The changes in the index are defined as a weighted average of the changes in the 90-day commercial paper rate and the changes in a trade-weighted Can$ exchange rate. Although the MCI was computed backward and onward from 1987, the Overnight Money Market Rate (available from 1975) is clearly a superior indicator of the Bank's policy stance | M1, M2plus# |
| *SWEDEN* | During the fixed-exchange rate regime, the overnight rate in the interbank market represented Riksbank's favourite instrument to keep the krona within the desired parity. Then, after the switch to the inflation-targeting regime, the Repo rate has become the operational instrument of the Bank. For the sake of homogeneity and continuity we use the Rate on 3-month Treasury Discount Notes (not available before 1982), which roughly shadows the behaviour of both marginal and Repo rates (Baumgartner et al., 1997). | M3 |
| *NEW ZEALAND* | The Rate on 90-day Bank Bills (not available before 1974) was our choice. Until March 1985, New Zealand has pursued an adjustable pegged exchange rate. "...the instrument since 1985 has been the quantity target for settlement balances held at the Reserve Bank. Settlement cash is used by commercial banks for end-of-day settlements with each other and the government. Should the banks run out of cash during the settlement period, further cash is available from the Reserve Bank by discounting Reserve Bank bills of short maturity at a penalty rate of 1.5% above market rates...Such an approach allows interest rates to move quickly, particularly when the change involves a politically unpopular increase in interest rates..." (Fischer, 1995, p.35) It is then understandable why banks prefer to act in the bank bills market, whose short-term interest rate tends to react rapidly to changes in policy intentions. | M1 |

* The Bundesbank announced targets for the growth of Central Bank Money until 1987, when it switched to M3, which we chose. The two move very closely together, apart from two episodes of divergence in 1988 and 1990-91. Despite the official target is announced in terms of base-money growth, the evidence points to Germany as to an "atypical" inflation targeter, who influences the money markets through changes in a day-to-day rate (Neumann and von Hagen, 1993; von Hagen, 1995; Bernanke and Mihov, 1998; Mishkin and Posen, 1997).

# Until 1982 the Bank of Canada was committed to target M1. It is now following closely also the behaviour of M2+ and a MCI, to obtain some indication about future inflation (Freedman, 1995).

| Test | GDP | Inflation '72-'97 | Inflation '80-'97 |
|---|---|---|---|
| | | **USA** | |
| **LC** | 0.548* | 0.543* | 1.762*** |
| **MeanF** | 7.946** | 10.713*** | 9.495*** |
| **SupF** | 23.682*** | 68.269*** | 48.400*** |
| | | **Germany** | |
| **LC** | 0.339 | 1.553*** | 0.651** |
| **MeanF** | 4.931 | 14.422*** | 8.195*** |
| **SupF** | 47.889*** | 25.059*** | 58.105*** |
| | | **Japan** | |
| **LC** | 0.193 | 0.743** | 0.998*** |
| **MeanF** | 4.186 | 5.913* | 7.987*** |
| **SupF** | 8.647 | 29.530*** | 28.035*** |
| | | **United Kingdom** | |
| **LC** | 0.514* | 0.624** | 2.049*** |
| **MeanF** | 13.150*** | 5.915** | 11.247*** |
| **SupF** | 43.787*** | 29.530*** | 48.371*** |
| | | **Canada** | |
| **LC** | 0.254 | 1.406*** | 1.383*** |
| **MeanF** | 5.449* | 12.803*** | 12.5*** |
| **SupF** | 24.662*** | 49.207*** | 55.908*** |
| | | **Sweden** | |
| **LC** | 0.334 | 1.137*** | 0.959*** |
| **MeanF** | 7.846** | 10.848*** | 24.192*** |
| **SupF** | 14.731* | 31.085*** | 73.784*** |
| | | **New Zealand** | |
| **LC** | 0.481* | 1.097*** | 0.484* |
| **MeanF** | 9.047*** | 21.772*** | 8.051*** |
| **SupF** | 26.849*** | 92.627*** | 27.016*** |

**Table 1 – Tests of parameter instability.** $L_c$, *MeanF*, *SupF* are defined as testing the null of stability against nonconstancy on the parameters of univariate autoregressive models for inflation (4-quarter change in CPI) and real GDP. Constants and linear time trends where included when relevant. *, **, and *** indicate significance of the relative F-statistic at the 10%, 5% and 1%, respectively (for tabulated critical values, see Hansen, 1992b).

| Sample/ Regressor | 1971Q4-1996Q3* | | 1980Q1-1996Q3* | | 1985Q1-1996Q3** | |
|---|---|---|---|---|---|---|
| Constant | 0.02213 (0.02348) | | 0.006149 (0.02366) | | 0.02422 (0.007616) | |
| Expected Inflation | 1.18 (0.4148) | | 1.81 (0.5315) | | 1.079 (0.2148) | |
| Output Gap | 1.572 (0.7553) | | 0.9438 (0.6183) | | 0.9266 (0.1387) | |
| Variable Addition Tests*** | money growth | 0.02210 [0.9781] | money growth | 0.04188 [0.9590] | money growth | 0.05796 [0.9438] |
| | exchange rate | 1.1505 [0.3211] | exchange rate | 0.36603 [0.6950] | exchange rate | 0.11224 [0.8941] |
| Summary Statistics | $R^2$ | 0.811253 | $R^2$ | 0.814444 | $R^2$ | 0.950583 |
| | $\sigma$ | 0.0151201 | $\sigma$ | 0.0165079 | $\sigma$ | 0.00440362 |
| | DW | 2.14 | DW | 2.19 | DW | 1.95 |
| | AR 1-5 $F(5, 90)$ | 5.0224 [0.0004] | AR 1-5 $F(5, 57)$ | 3.42 [0.0090] | AR 1-5 $F(5, 40)$ | 0.76539 [0.5802] |
| | ARCH 4 $F(4, 87)$ | 15 [0.0000] | ARCH 4 $F(4, 54)$ | 17.675 [0.0000] | ARCH 4 $F(4, 37)$ | 0.48356 [0.7476] |
| | Normality $\chi^2(2)$ | 63.105 [0.0000] | Normality $\chi^2(2)$ | 65.253 [0.0000] | Normality $\chi^2(2)$ | 2.9124 [0.2331] |
| | RESET $F(1, 94)$ | 0.38632 [0.5357] | RESET $F(1, 61)$ | 1.6571 [0.2029] | RESET $F(1, 44)$ | 8.28e-007 [0.999] |

**Table 2 – USA. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and two lags of the dependent variable.*

*** Derived from RLS regression of the interest rate on a constant, 2-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

****We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in a lagged trade-weighted index of effective exchange rate were tested by a F-version of the Wald test on the baseline model augmented of each new variable. P-values in brackets.*

*Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P-values in brackets.*

| Sample/ Regressor | 1971Q4-1996Q3 | | 1982Q1-1996Q3 | |
|---|---|---|---|---|
| | *Baseline** | *Adding Fed Funds Rate*** | *Baseline** | *Adding Fed Funds Rate*** |
| **Constant** | 0.03263<br>(0.01081) | -0.01123<br>(0.01276) | -7.478<br>(2.538) | -0.02378<br>(0.01929) |
| **Expected Inflation** | 0.6292<br>(0.1894) | 0.4389<br>(0.09508) | 1.872<br>(0.6319) | 1.821<br>(0.5251) |
| **Output Gap** | 0.791<br>(0.5732) | -0.03604<br>(0.3074) | 1.22<br>(0.7698) | 0.8286<br>(0.5059) |
| **Fed Funds Rate** | | 0.6364<br>(0.1346) | | 0.5548<br>(0.3044) |
| Variable Addition Tests*** | money growth 1.3935 [0.2533]<br>exchange rate 2.4543 [0.0684] | | money growth 0.29414 [0.7464]<br>exchange rate 0.6417 [0.5917] | |
| Summary Statistics | $R^2$ 0.95232<br>$\sigma$ 0.00639043<br>DW 2.24<br>AR 1-5 $F(5, 90)$ 3.1778 [0.0109]<br>ARCH 4 $F(4, 87)$ 2.1103 [0.0863]<br>Normality $\chi^2(2)$ 25.49 [0.0000]<br>RESET $F(1, 94)$ 4.189 [0.0435] | $R^2$ 0.960445<br>$\sigma$ 0.00585134<br>DW 2.17<br>AR 1-5 $F(5, 89)$ 1.447 [0.2154]<br>ARCH 4 $F(4, 86)$ 2.4588 [0.0514]<br>Normality $\chi^2(2)$ 9.3293 [0.0094]<br>RESET $F(1, 93)$ 1.7085 [0.1944] | $R^2$ 0.971225<br>$\sigma$ 0.00412458<br>DW 1.90<br>AR 1-5 $F(5, 53)$ 0.2559 [0.9350]<br>ARCH 4 $F(4, 50)$ 8.2479 [0.0000]<br>Normality $\chi^2(2)$ 5.1598 [0.0758]<br>RESET $F(1, 57)$ 3.2564 [0.0764] | $R^2$ 0.969527<br>$\sigma$ 0.00398603<br>DW 2.05<br>AR 1-5 $F(5, 48)$ 0.46784 [0.7983]<br>ARCH 4 $F(4, 45)$ 3.5538 [0.0133]<br>Normality $\chi^2(2)$ 6.9901 [0.0303]<br>RESET $F(1, 52)$ 0.092264[0.7625] |

**Table 3 – Japan. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and two lags of the dependent variable.*

***As for the note above, but now with one lag of the Fed Funds Rate on the RHS.*

****We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in the current, one- and twice-lagged trade weighted exchange rate were tested by a F-version of the Wald test. P-values in brackets.*

*Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P-values in brackets.*

| *Sample/ Regressor* | *1970Q3-1996Q4* | | *1980Q1-1996Q4* | |
|---|---|---|---|---|
| | *Baseline** | *Adding Fed Funds Rate*** | *Baseline** | *Adding Fed Funds Rate*** |
| **Constant** | 0.01284<br>(0.01498) | -0.002516<br>(0.01731) | 0.02088<br>(0.00818) | 0.007248<br>(0.01178) |
| **Expected Inflation** | 1.416<br>(0.3835) | 1.174<br>(0.3255) | 1.494<br>(0.2434) | 1.373<br>(0.2602) |
| **Output Gap** | 0.9186<br>(0.4643) | 0.8073<br>(0.3999) | 0.4848<br>(0.2351) | 0.6077<br>(0.263) |
| **Fed Funds Rate** | | 0.3144<br>(0.1702) | | 0.21<br>(0.1175) |
| *Variable Addition Tests**** | money growth 1.2556 [0.2895]<br>exchange rate 1.2552 [0.2948] | | money growth 0.56493 [0.5713]<br>exchange rate 0.71692 [0.5457] | |
| *Summary Statistics* | $R^2$ 0.870249<br>$\sigma$ 0.009158<br>DW 1.73<br>AR 1-5 F(5, 97) 1.2563 [0.2891]<br>ARCH 4 F(4, 94) 3.3586 [0.0129]<br>Normality $\chi^2(2)$ 55.873 [0.0000]<br>RESET F(1,101) 0.4661 [0.4963] | $R^2$ 0.8908<br>$\sigma$ 0.008485<br>DW 1.80<br>AR 1-5 F(5, 95) 1.4545 [0.2121]<br>ARCH 4 F(4, 92) 2.3624 [0.0588]<br>Normality $\chi^2(2)$ 65.731 [0.0000]<br>RESET F(1, 99) 1.2401 [0.2682] | $R^2$ 0.949219<br>$\sigma$ 0.0055634<br>DW 1.69<br>AR 1-5 F(5, 59) 0.7215 [0.6100]<br>ARCH 4 F(4, 56) 0.4445 [0.7759]<br>Normality $\chi^2(2)$ 11.187 [0.0037]<br>RESET F(1, 63) 0.0168 [0.8971] | $R^2$ 0.963906<br>$\sigma$ 0.00476544<br>DW 1.60<br>AR 1-5 F(5, 57) 2.0422 [0.0863]<br>ARCH 4 F(4, 54) 1.4412 [0.2330]<br>Normality $\chi^2(2)$ 0.3125 [0.8554]<br>RESET F(1, 61) 0.0301 [0.8627] |

**Table 4. Germany. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and one lag of the dependent variable.*
***As for the note above, but now with two lags of the Fed Funds Rate on the RHS.*
****We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in the current and lagged exchange rate vis-a-vis the US$ were tested by a F-version of the Wald test. P-values in brackets.*
*Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P-values in brackets.*

| Sample/ Regressor | 1975Q3-1996Q3 | | | | 1980Q1-1996Q3 | | 1985Q1-1996Q3 | |
|---|---|---|---|---|---|---|---|---|
| | Baseline* | | Adding Exchange Rate** | | Baseline* | | Baseline*** | |
| Constant | 0.04062 (0.02056) | | 0.02662 (0.02697) | | 0.0393 (0.0187) | | 0.02367 (0.01275) | |
| Expected Inflation | 0.9088 (0.2821) | | 1.117 (0.3775) | | 1.087 (0.322) | | 1.403 (0.2831) | |
| Output Gap | 0.8017 (0.3672) | | 0.9981 (0.4593) | | 0.6779 (0.3344) | | 0.6388 (0.1987) | |
| Exchange Rate | | | -0.6601 (0.3609) | | | | | |
| Variable Addition Tests**** | M4 growth | 2.2368 [0.1135] | | | money growth | 0.84828 [0.4331] | moneygrowth | 4.7366 [0.0141] |
| | M1 growth | 2.4012 [0.0972] | | | | | | |
| | exchange rate | 2.5803 [0.0595] | | | exchange rate | 1.565 [0.2072] | exchange rate | 2.0764 [0.1186] |
| | German rate | 3.598 [0.0173] | | | German rate | 1.5552 [0.2096] | German rate | 3.7459 [0.0184] |
| Summary Statistics | $R^2$ | 0.867889 | $R^2$ | 0.878873 | $R^2$ | 0.904109 | $R^2$ | 0.93223 |
| | $\sigma$ | 0.011124 | $\sigma$ | 0.0107179 | $\sigma$ | 0.00982599 | $\sigma$ | 0.00844755 |
| | DW | 1.67 | DW | 1.58 | DW | 1.74 | DW | 1.27 |
| | AR 1-5 $F(5, 76)$ | 0.81065 [0.5457] | AR 1-5 $F(5, 75)$ | 1.1024 [0.3664] | AR 1-5 $F(5, 58)$ | 0.83746 [0.5286] | AR 1-5 $F(5, 38)$ | 2.3838 [0.0563] |
| | ARCH 4 $F(4, 73)$ | 1.6248 [0.1772] | ARCH 4 $F(4, 72)$ | 0.98115 [0.4233] | ARCH 4 $F(4, 55)$ | 1.2151 [0.3150] | ARCH 4 $F(4, 35)$ | 1.8764 [0.1365] |
| | Normality $\chi^2(2)$ | 7.1475 [0.0281] | Normality $\chi^2(2)$ | 5.4778 [0.0646] | Normality $\chi^2(2)$ | 8.6269 [0.0134] | Normality $\chi^2(2)$ | 12.864 [0.0016] |
| | RESET $F(1, 80)$ | 0.0955 [0.7581] | RESET $F(1, 79)$ | 0.18972 [0.6643] | RESET $F(1, 62)$ | 1.2575 [0.2664] | RESET $F(1, 42)$ | 2.4368 [0.1260] |

**Table 5. United Kingdom. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

***Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and the current trade-weighted index of effective exchange rate.*

****Derived from RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

*****We tested for the addition of other regressors. Zero restrictions on lagged money growth (both M4 and M1), changes in the current and lagged trade-weighted index of effective exchange rate and current and lagged 3-month German FIBOR were tested by a F-version of the Wald test. P-values in brackets.*

| Sample/ Regressor | 1975Q3-1996Q2 | 1975Q3-1996Q2 | 1982Q1-1996Q2 |
|---|---|---|---|
| | Baseline* | Adding Federal Funds Rate** | Adding Federal Funds Rate*** |
| **Constant** | 0.04624<br>(0.01912) | 0.01807<br>(0.00929) | 0.02178<br>(0.007424) |
| **Expected Inflation** | 1.036<br>(0.3289) | 0.007358<br>(0.1505) | 0.2022<br>(0.2541) |
| **Output Gap** | 0.6367<br>(0.5969) | 0.3783<br>(0.209) | 0.224<br>(0.1501) |
| **Fed Funds Rate** | | 1.009<br>(0.1523) | 0.8696<br>(0.161) |
| Variable Addition Tests**** | money growth 1.168 [0.3164]<br>exchange rate 0.54758 [0.6513] | | exchange rate 3.2878 [0.0281] |
| Summary Statistics | $R^2$ 0.740868<br>$\sigma$ 0.0184464<br>DW 1.98<br>AR 1-4 F( 4, 72) 0.10383 [0.9808]<br>ARCH 4 F( 4, 68) 6.6823 [0.0001]<br>Normality $\chi^2(2)$ 20.536 [0.0000]<br>RESET F( 1, 75) 0.02468[0.8756] | $R^2$ 0.818641<br>$\sigma$ 0.0152328<br>DW 2.32<br>AR 1-4 F( 4, 74) 6.8395 [0.0001]<br>ARCH 4 F( 4, 70) 7.2532 [0.0001]<br>Normality $\chi^2(2)$ 4.1906 [0.1230]<br>RESET F( 1, 77) 0.3653[0.5473] | $R^2$ 0.907322<br>$\sigma$ 0.00879038<br>DW 1.96<br>AR 1-4 F( 4, 47) 0.37725 [0.8237]<br>ARCH 4 F( 4, 43) 1.9921 [0.1128]<br>Normality $\chi^2(2)$ 0.75685 [0.6849]<br>RESET F( 1, 50) 0.14571 [0.7043] |

**Table 6. Canada. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap and two lags of the dependent variable.*

***Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, output gap, two lags of the dependent variable and the current Fed Funds Rate.*

****Derived from RLS regression of the interest rate on a constant, 4-quarter ahead expected inflation, current and lagged output gap, one lag of the dependent variable and current Fed Funds Rate.*

*****We tested for the addition of other regressors. Zero restrictions on lagged money growth and changes in the current and lagged trade-weighted index of effective exchange rate were tested by a F-version of the Wald test. P-values in brackets.*

| *Sample/ Regressor* | *1982Q3-1997Q2* | | | |
|---|---|---|---|---|
| | *Baseline** | | *Adding ERM dummy*** | |
| ***Constant*** | 0.06397 (0.01735) | | 0.06238 (0.00912) | |
| ***Expected Inflation*** | 0.6763 (0.2811) | | 0.7111 (0.1471) | |
| ***Output Gap*** | 0.4803 (0.4653) | | 0.4461 (0.2389) | |
| ***ERM dummy*** | | | 0.08641 (0.02487) | |
| *Variable Addition Tests**** | | | money growth | 2.3225 [0.1079] |
| | | | exchange rate | 1.6609 [0.1868] |
| | | | German rate | 2.5539 [0.0654] |
| *Summary Statistics* | $R^2$ | 0.826297 | $R^2$ | 0.852995 |
| | $\sigma$ | 0.0117457 | $\sigma$ | 0.01091 |
| | DW | 1.66 | DW | 1.42 |
| | AR 1- 4 $F(4, 52)$ | 2.2564 [0.0755] | AR 1- 4 $F(4, 51)$ | 3.238 [0.018] |
| | ARCH 4 $F(4, 48)$ | 0.26744 [0.8975] | ARCH 4 $F(4, 47)$ | 2.653 [0.044] |
| | Normality $\chi^2(2)$ | 10.808 [0.0045] | Normality $\chi^2(2)$ | 2.238 [0.327] |
| | RESET $F(1, 55)$ | 3.463 [0.0681] | RESET $F(1, 54)$ | 3.48 [0.0676] |

**Table 7. Sweden. Solved Static Long-run Equations**

**Derived from RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

***Derived from RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and a dummy variable assuming value one in the third and fourth quarter on 1992 and zero elsewhere, and one lag of the dependent variable.*

****We tested for the addition of other regressors. Zero restrictions on lagged money growth, changes in the current and lagged trade-weighted index of effective exchange rate and current and lagged 3-month German FIBOR were tested by a F-version of the Wald test. P-values in brackets.*

| Sample/ Regressor | 1982Q4-1997Q2* | | 1982Q4-1997Q1** | |
|---|---|---|---|---|
| **Constant** | 0.06188 (0.00819) | | 0.06278 (0.00995) | |
| **Expected Inflation** | 1.105 (0.112) | | 1.106 (0.1315) | |
| **Output Gap** | 0.01263 (0.1858) | | -0.3916 (0.1965) | |
| Variable Addition Tests*** | Money growth | 2.5969 [0.0844] | | |
| | Exchange rate | 1.8725 [0.1457] | | |
| Summary Statistics | $R^2$ | 0.933781 | $R^2$ | 0.940185 |
| | $\sigma$ | 0.0143532 | $\sigma$ | 0.0138594 |
| | DW | 1.93 | DW | 1.76 |
| | AR 1-4 $F(4, 51)$ | 0.21108 [0.9311] | AR 1-4 $F(4, 46)$ | 0.58269 [0.6767] |
| | ARCH 4 $F(4, 47)$ | 5.3048 [0.0013] | ARCH 4 $F(4, 42)$ | 2.0449 [0.1054] |
| | Normality $\chi^2(2)$ | 1.6867 [0.4303] | Normality $\chi^2(2)$ | 11.539 [0.0031] |
| | RESET $F(1, 54)$ | 1.6545 [0.2038] | RESET $F(1, 49)$ | 0.11753 [0.7332] |

**Table 8. New Zealand. Solved Static Long-run Equations**

**Derived from a RLS regression of the interest rate on a constant, one-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

***Derived from a RLS regression of the interest rate on a constant, two-quarter ahead expected inflation, output gap and one lag of the dependent variable.*

****We tested for the addition of other regressors. Zero restrictions on lagged money growth, changes in the current and lagged trade-weighted index of effective exchange rate were tested by a F-version of the Wald test. P-values in brackets.*

*Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P-values in brackets.*

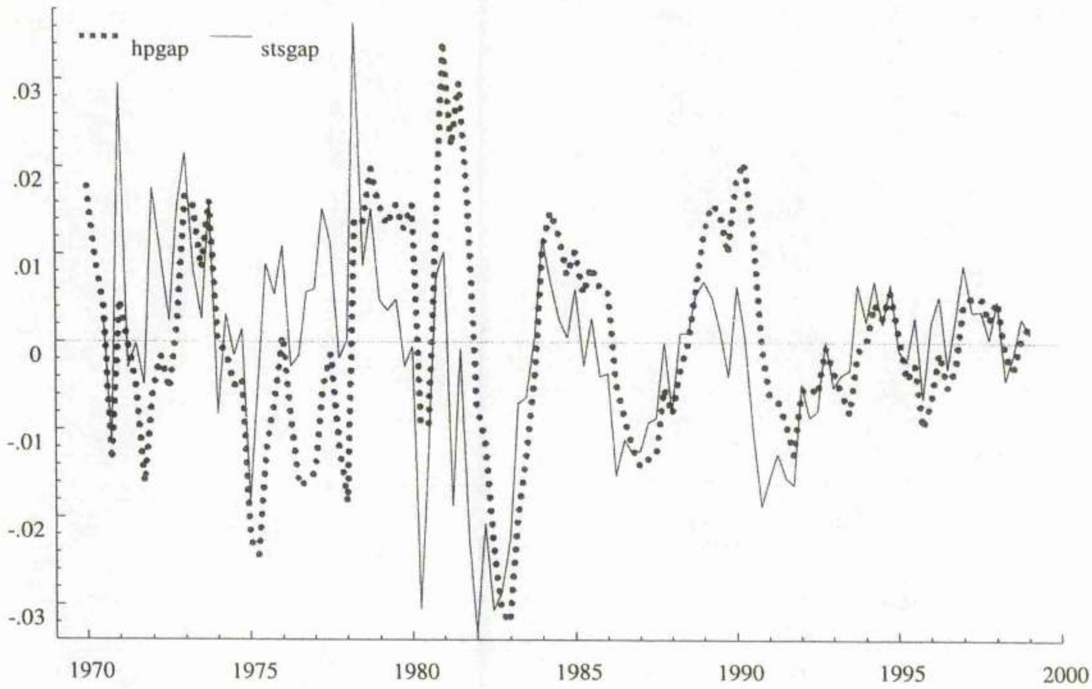


**Figure 1. USA, 1970Q1-1998Q4.** Hodrick-Prescott and Filtered STS measures of the Output Gap

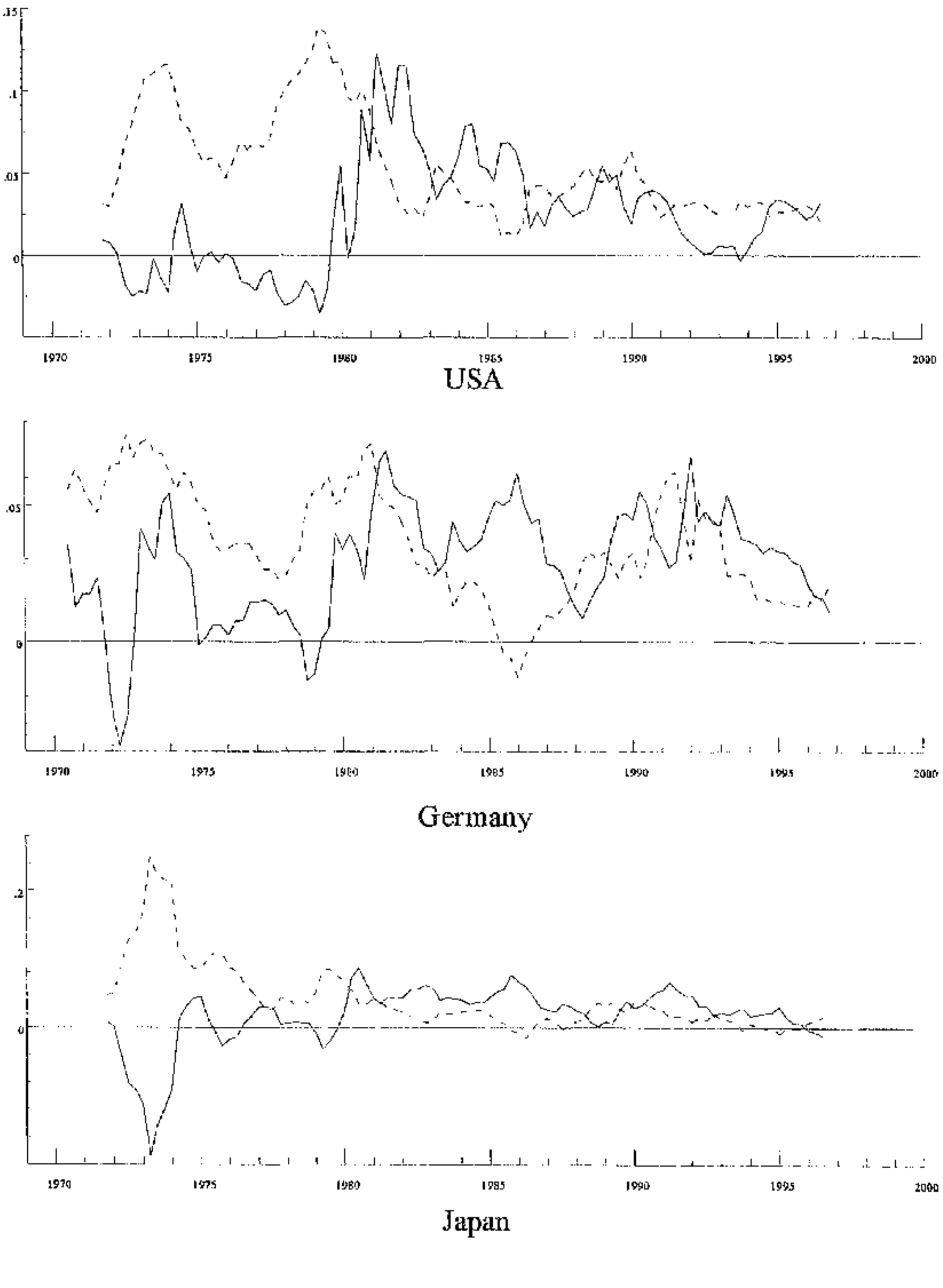


**Figure 2** - USA, Germany, Japan. *Ex ante* real interest rates (solid lines) and (4-quarter ahead) expected inflation (dotted lines)

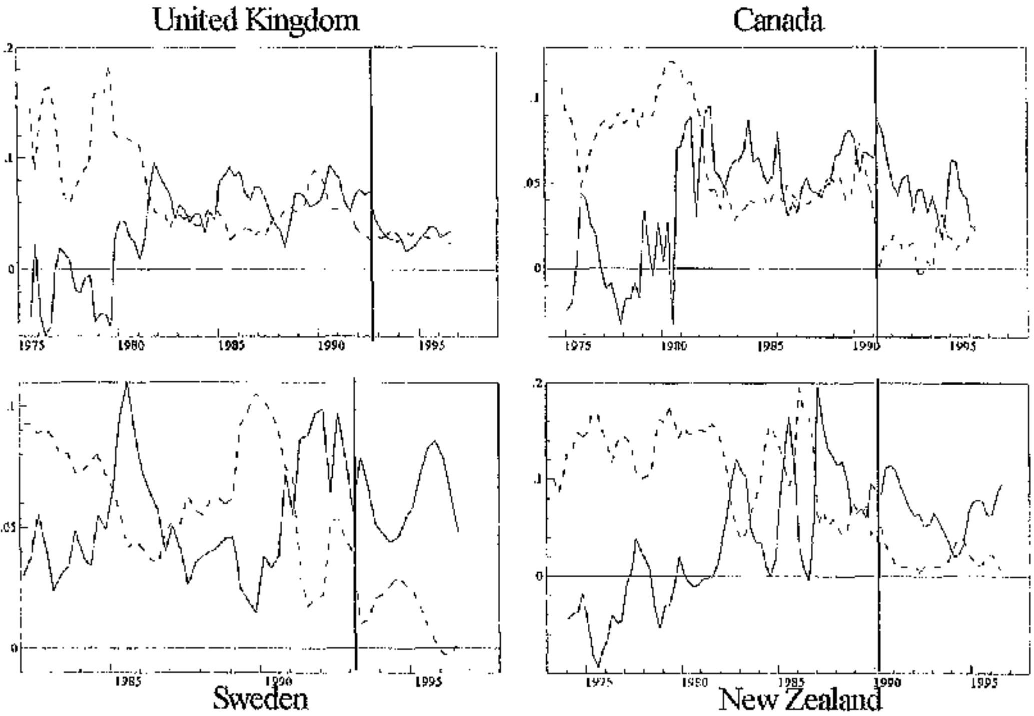


**Figure 3** - United Kingdom, Canada, Sweden, New Zealand. *Ex ante* real interest rates (solid lines) and (4-quarter ahead) expected inflation (dotted lines). The vertical lines represent the announcement of inflation targets

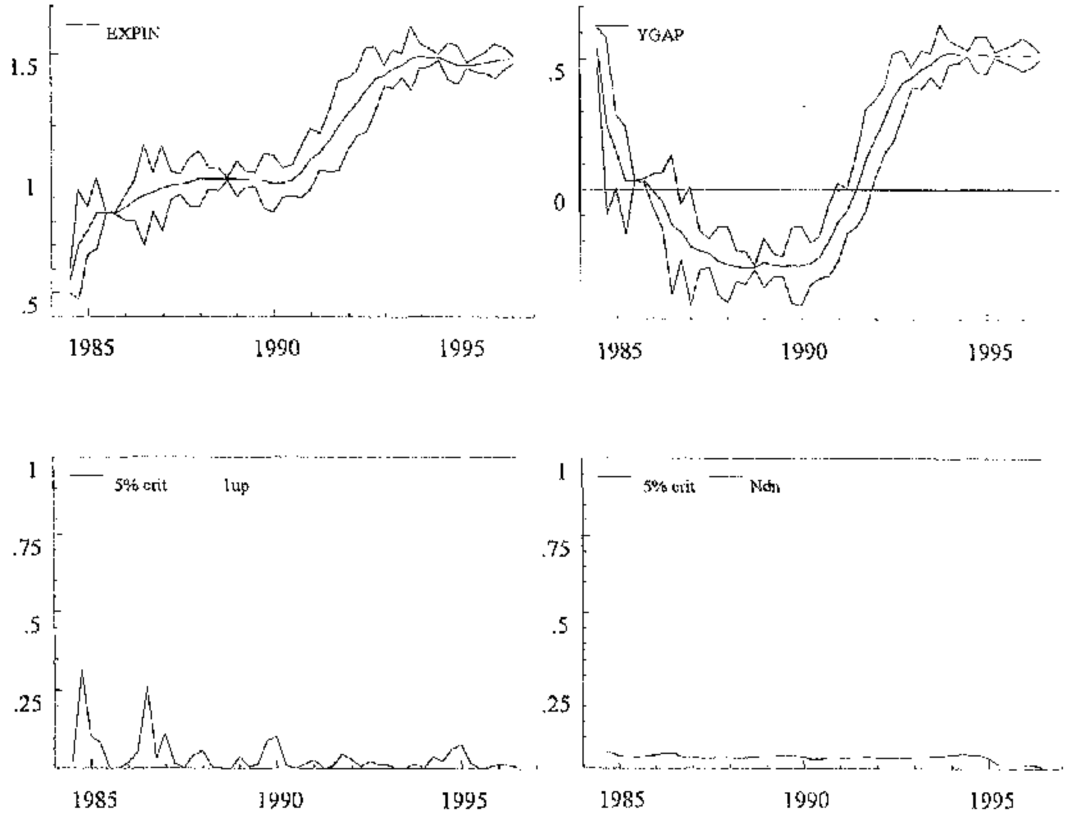


**Figure 4. USA**, 1980(1)-1996(3). Recursive coefficients and standard error bands; 1-step up, N-step down Chow tests (5%)

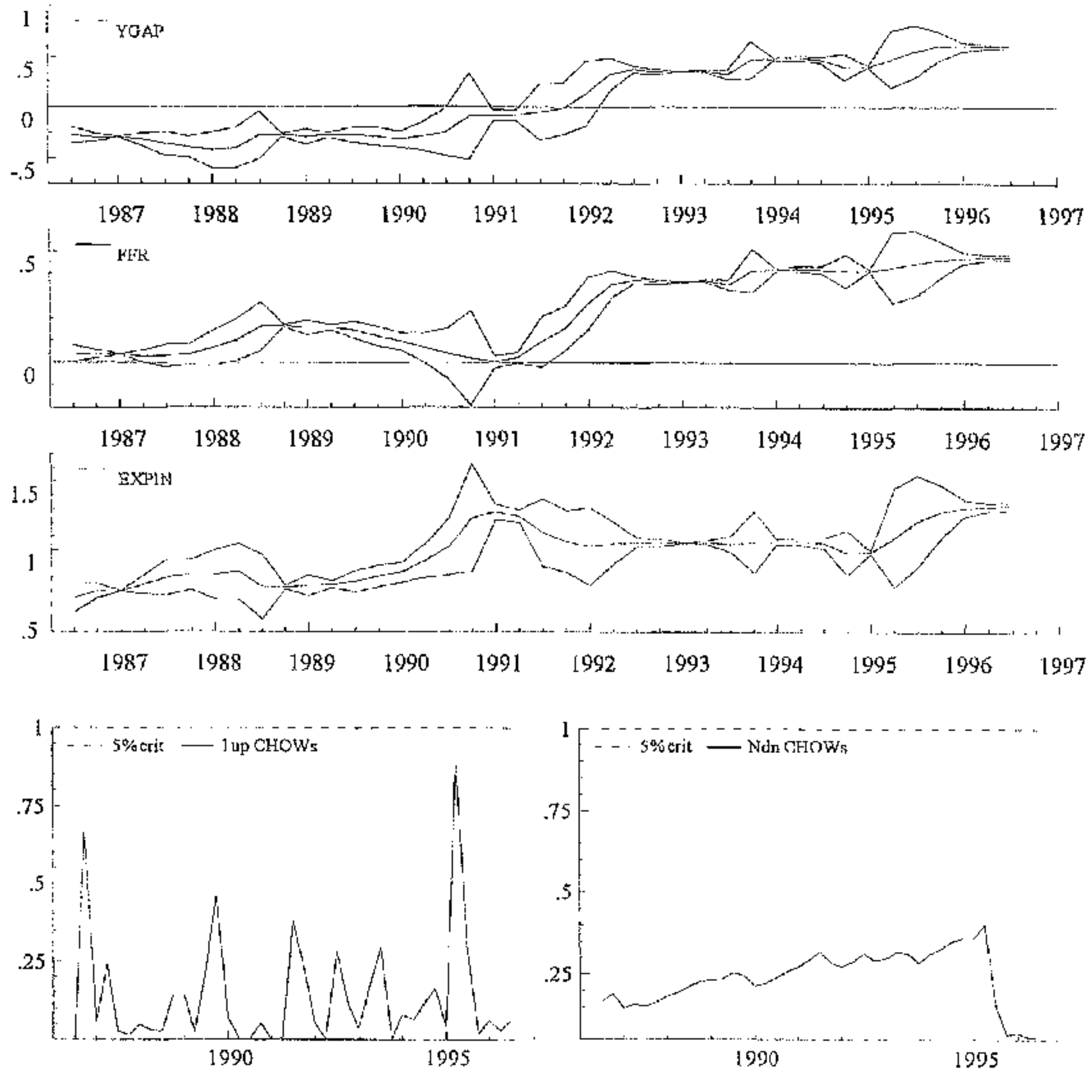


**Figure 5. Japan**, 1982(1)-1996(3). Recursive coefficients and standard error bands; 1-step up, N-step down Chow tests (5%)

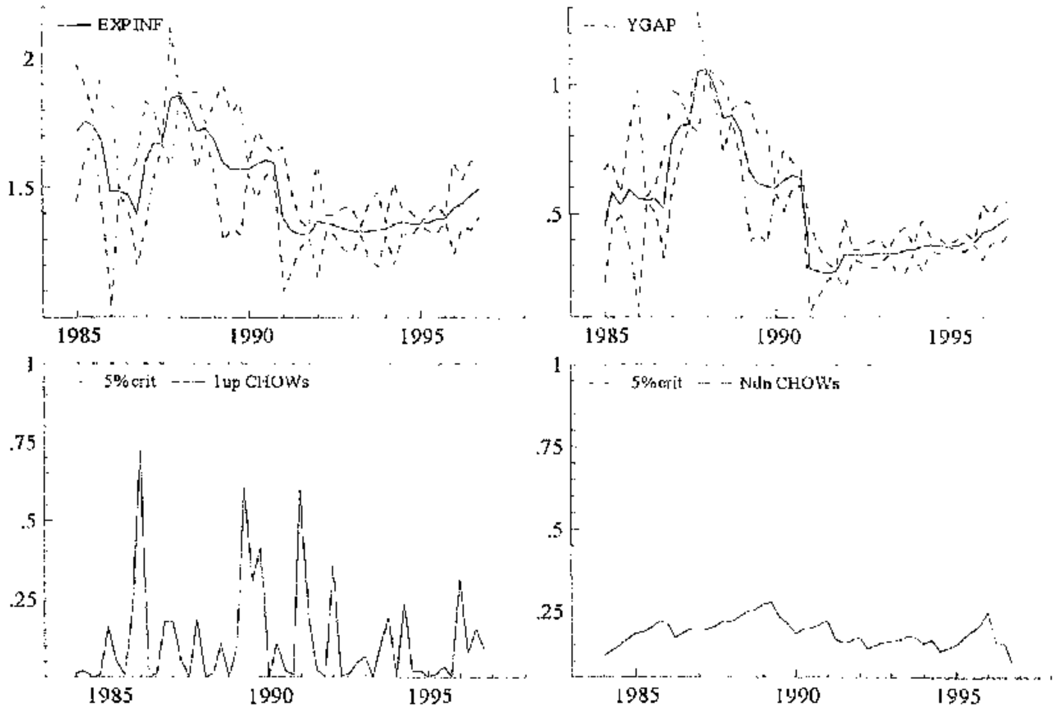


**Figure 6. Germany,** 1980(1)-1996(4). Recursive coefficients and standard error bands; 1-step, N-step up Chow tests (5%)

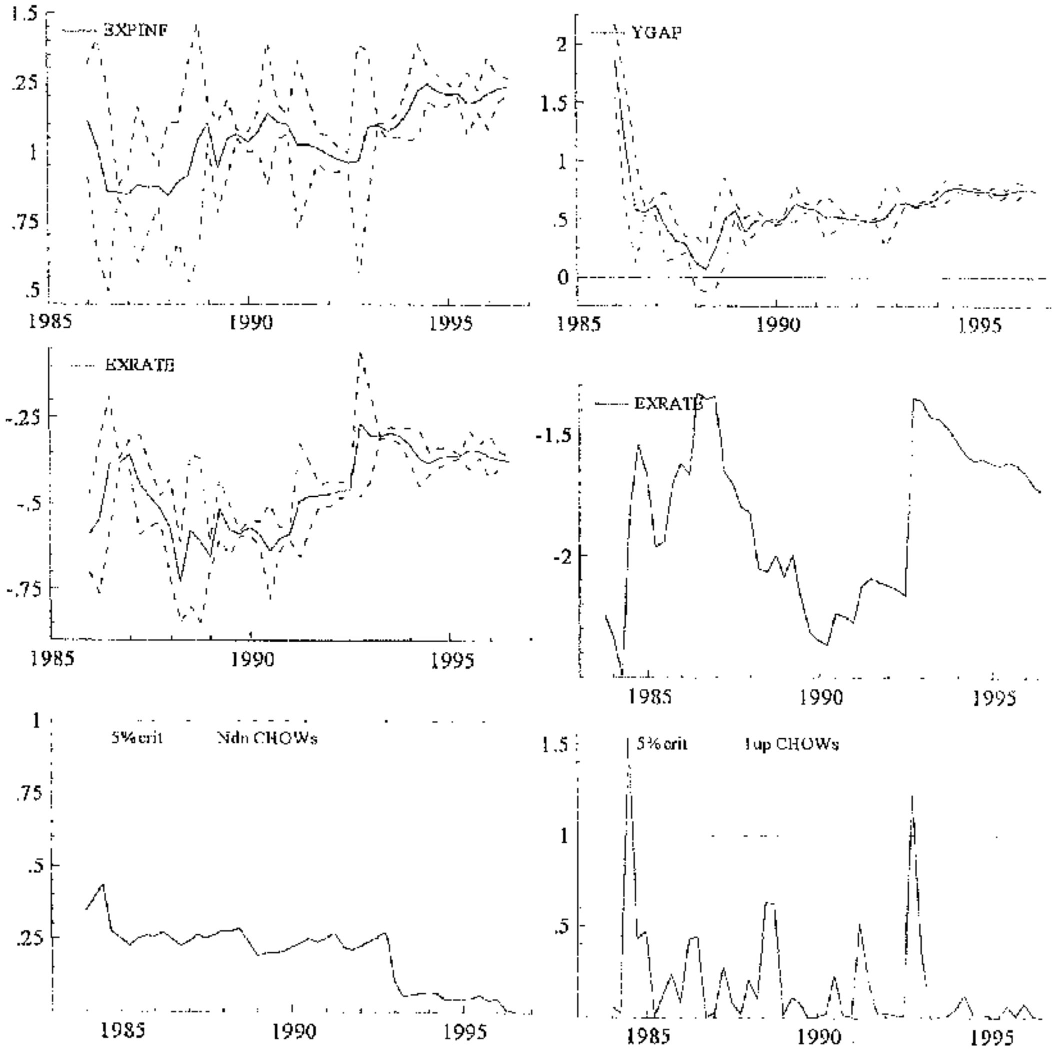


**Figure 7. United Kingdom**, 1980(1)-1996(3). Recursive coefficients and standard error bands; t-ratio for the exchange rate (short-run) coefficient; 1-step up, N-step up Chow tests (5%)

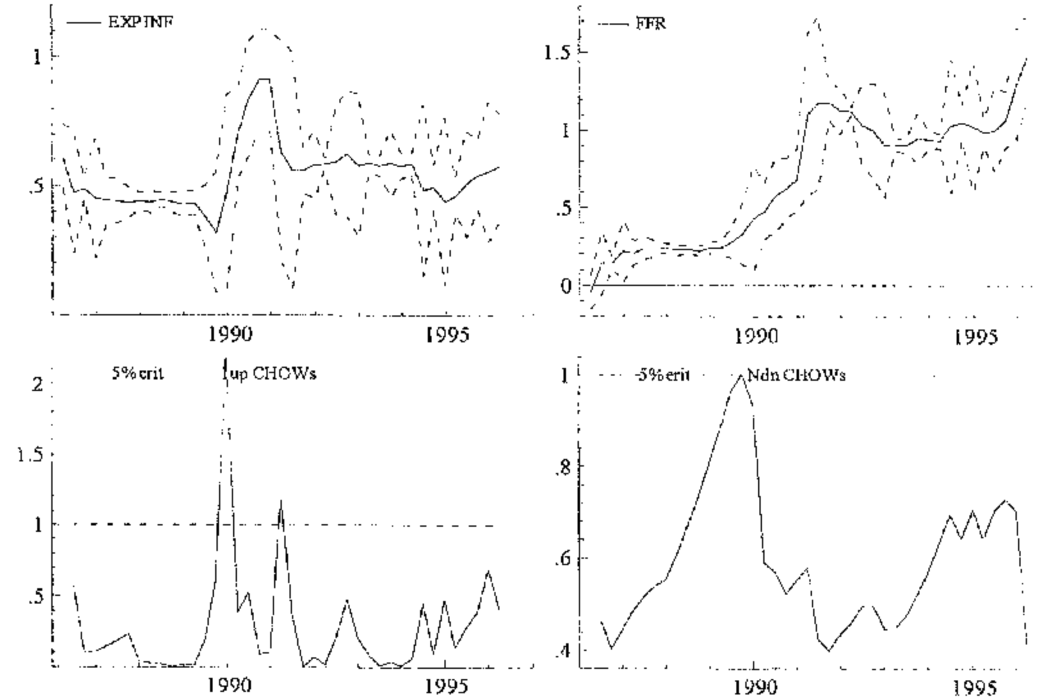


**Figure 8. Canada,** 1982(2)-1996(2). Recursive coefficients and standard error bands; 1-step up, N-step down Chow tests (5%)

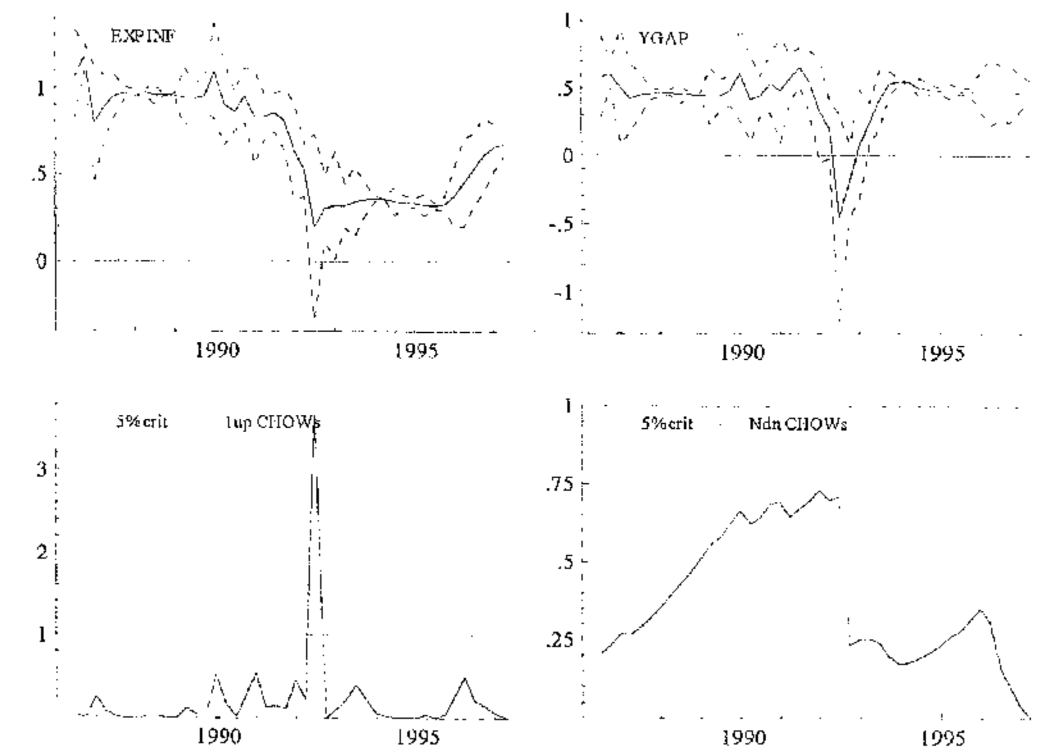


**Figure 9. Sweden,** 1982(2)-1997(2). Recursive coefficients and standard error bands; 1-step up, N-step down Chow tests (5%)

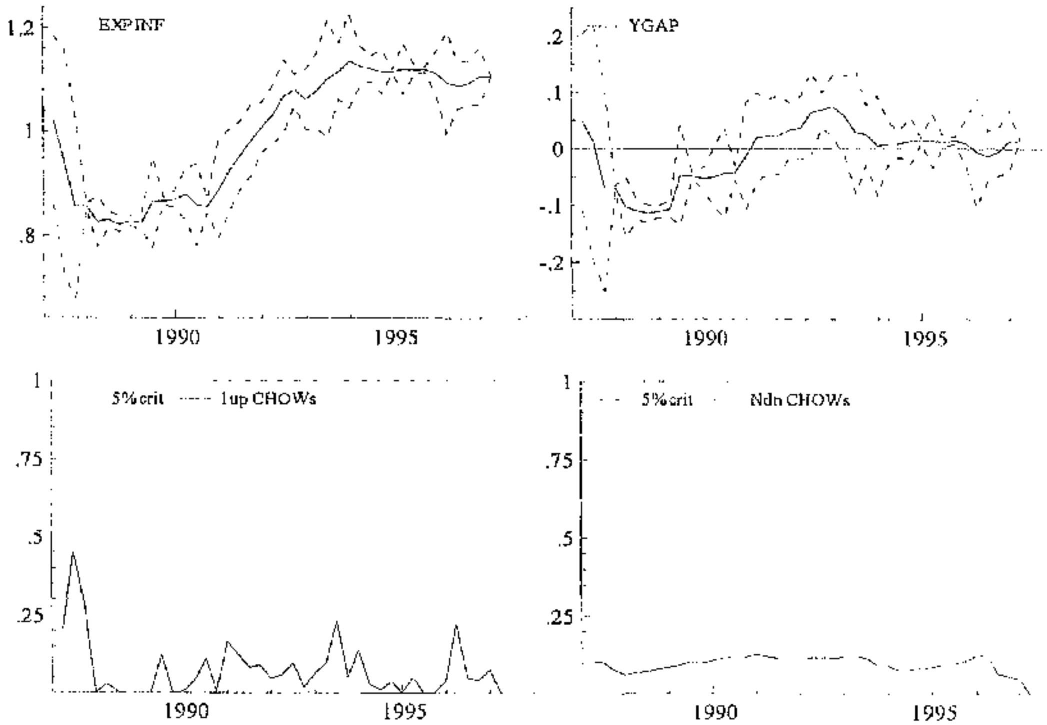


**Figure 10. New Zealand**, 1982(4)-1997(2). Recursive coefficients and standard error bands; 1-step up, N-step down Chow tests (5%)

# Interest Rate Rules and Policy Credibility from ERM to EMU

# Chapter 2

*"...But how do you run a common currency without a common government? Europe has some experience with this sort of thing. For almost two decades - since the formation of the European Monetary System in 1979 - most European nations have committed themselves to maintaining fixed exchange rates between their currencies, which basically means adopting a common monetary policy. And while there have been occasional flareups in the arrangement - a last-gasp attempt by the French to follow their own path back in 1982, and a wave of speculative attacks that pushed Britain out of the system a decade later - the EMS has proved surprisingly durable. How did Europe manage to follow a common monetary policy? There was a bit of neatly calculated hypocrisy. Athough the EMS was in principle a symmetric system, with all countries treated equally, in practice it was tacitly run as a German hegemony: the Bundesbank set interest rates as it pleased, and other central banks then did whatever was necessary to keep their currencies pegged to the Deutsche mark. This arrangement allowed the system to meet two seemingly irreconcilable demands: the insistence of Germans, who still remember both the hyperinflation of 1923 and the economic miracle that followed the introduction of a new, stable currency in 1948, that their beloved Bundesbank keep its hand firmly on the monetary tiller; and the political imperative that any European institution must look like an association of equals, not a new, um, Reich. The Europeans, they are a subtle race..."*

P. Krugman, *Fortune*, December 1998

## 1. Introduction

In the previous chapter, we studied how changes in monetary policy institutions influence the way interest rate policies react to expected inflation and the business cycle. We saw that, in some instances, greater central bank independence and the introduction of inflation targets do not seem to have radically altered the way in which authorities react to changes in the final objectives of monetary policy. The latter were identified using a simple policy rule expressed in terms of deviations of expected inflation from some pre-set target, and a measure of the output gap.

In this chapter, we investigate the same issue, this time with reference to four former EMS countries. More precisely, we study, from a perspective similar to that in the previous chapter, the effects of the most remarkable institutional change monetary policy has undergone in modern times, namely, the process of monetary unification in Europe. Whether such process will be successful or not is a complex and perhaps unanswerable question. Instead, we wish to evaluate the path followed by monetary policies throughout Europe in the process of monetary convergence towards EMU. More in detail, we ask ourselves whether and how the nominal convergence achieved amongst the former EMS countries was affected by the existence of the Exchange Rate Mechanism, and by other well-known constraints on national monetary policies. Apart from the existence of exchange rate bands, we attempt a broad assessment of the way in which the required convergence in budget positions across countries has affected the response of national monetary authorities to final domestic objectives.

This broad-based aim suggested us to estimate interest rate reaction functions - similarly to what we did for the G-3 and inflation-targeting economies- for four key European countries: France, Italy, Ireland and Belgium. Our estimation sample -1980Q1-1997Q2- covers the whole period spanned by the EMS, and ends when financial markets started to be persuaded as to the real outcome of the Stage Two of EMU, i.e., whether the euro would have really be put in place as planned. We recall that the major doubts about the start of Stage Three of EMU concerned the soundness of recent efforts of fiscal consolidation produced by some countries (De Grauwe, 1997; Obstfeld, 1998; Dornbusch, Favero and Giavazzi, 1998; Eichengreen and Wyplosz, 1998). It is then interesting to understand the extent to which such efforts and those concerns influenced the course of national interest rate policies on the road to EMU. Once interest rate reaction functions for the above mentioned countries are estimated, we evaluate the stability of estimated equations and preference parameters attached to domestic objectives and relevant additional regressors. This allows us to draw some conclusions about the convergence process amongst these countries and Germany.

The question, besides an eminently historical interest, would certainly help understanding how ESCB's monetary policies will be drafted in the coming years. We wish to provide an approximate assessment of the costs and the adjustments in policy preferences that accompanied the process of monetary convergence throughout Europe. Such process involved countries where initial monetary conditions and policy credibility were very similar (Belgium), relatively similar (France), or relatively different (Italy and Ireland) from those prevailing in Germany. A better knowledge of the individual trade-offs faced by the monetary authorities of these countries would surely reveal some clue about how national issues will be evaluated in the current and future issues decision-making process of the ESCB.

The literature developed during the latter part of the 1980s provided an intuitive and flexible framework for studying the EMS (Giavazzi and Giovannini, 1989; Fratianni and von Hagen, 1992). The main theoretical motivation devised for the existence of the

EMS was that, *if* credible, it could have provided an effective instrument to bring monetary discipline to relatively inflation-prone countries like the ones we chose. In fact, conventional accounts of monetary policy events in Europe tend to argue that, since the second half of 1980s, a stronger exchange rate commitment helped in bringing down inflation expectations in many European countries (Caporale and Pittis, 1993)[1]. In other words, the exchange rate agreement apparently forced national policymakers to pursue more restrictive monetary policies than those that might have been followed in the absence of such agreements. The fall in inflation, and the following strong convergence in nominal terms achieved by the start of Stage Three of EMU, is then to be attributed to the ever stronger external constraints binding national authorities.

In reality, given the long history of realignments and parity adjustments, especially in the early years of its existence, one can safely argue that the ERM was a classical example of partially credible target zone regime. In other words, price stability and pan-European nominal convergence were achieved only after a long and uncertain process, in which the simple ERM membership was not sufficient to guarantee the outcome.

However, one can argue that the ERM, along with other factors, did have a "wheel-greasing" role in forcing such process. The well-known imported credibility approach to the EMS, stemming from Giavazzi and Pagano's (1988) work (see also Coles and Philippopoulos, 1997), argues that EMS membership was indeed crucial in bringing in potentially large credibility gains to national monetary authorities. By attaching extra penalties to departures of interest rate policies from an anti-inflationary stance, the ERM of the EMS made financial markets, and the public in general, more aware of the new trade-off faced by monetary authorities. This way, the exchange rate agreement provided low-credibility policymakers with some commitment technology, gradually shrinking the inefficiencies commonly associated with monetary policy equilibria in the absence of commitment (Persson and Tabellini, 1999). Last, but not least, severe ceilings on the use of national fiscal policies were enacted by 1992's Maastricht Treaty, adding further grounds on which the conduct of national policymakers were to be evaluated in EMU's perspective. In addition, throughout the latter half of 1980s, capital controls were progressively dismantled, financial markets became more integrated on an international

[1] However, inflation fell along a very similar pattern in the majority of OECD countries outside the EMS.

level, and increasing product market integration was overall promoted by EC-EU institutions.

While the exact impact of those and other changes is still difficult to ascertain, it is relatively safe to argue that they gradually made exchange rate realignments more costly and less effective in boosting competitiveness and growth. One can also hypothesise that such changing macroeconomic environment gradually tilted the balance between the benefits from realignments, and the credibility gains brought about by a more "disciplined" conduct of monetary policy, in favour of the latter.

It is then interesting to understand how the convergence of national Central Banks' -or the public's- preferences towards Bundesbank's anti-inflationary attitude[2] took place. Similarly, given the relevance attributed by the Maastricht Treaty and the subsequent Stability and Growth Pact, to fiscal consolidation, one might want to evaluate how such additional constraint affected Member States' macroeconomic conditions on their road to EMU.

In this chapter, we provide some empirical evidence supporting the view that the historical path followed by monetary policies in the former Members of EMS to achieve nominal convergence with Germany was not uniform. We show that the process itself bore significant shifts to the way in which monetary policy authorities were reacting to domestic objectives. This appears to have taken place with differentiated timings in Belgium, France, Ireland, and Italy. In addition, we allow our theoretical model to take into account the possibility that the credibility of the fiscal stance explicitly affected interest rate policies adopted in the EMS countries. With imperfect credibility, an unsustainable fiscal position in principle may induce markets to believe that the central bank will need to loosen its anti-inflationary stance (and, ultimately, the country's exchange rate commitment) in the future. This line of reasoning has been recently revived in the debate on the fiscal theory of price level determination (Woodford, 1995; Canzoneri and Diba, 1996; Cochrane, 2000).

Following Favero, Giavazzi and Spaventa (1997), we proxied market's perception of the exchange rate risk by an adjusted measure of the long-term yield differential between each country and Germany. If unbalanced fiscal policies were to affect market's perception of the probability of loose interest rate policies in the future, the optimal policy rule would directly target such perception. By consequence, and with reference to the policy rules analysed in the preceding chapter, the policy instrument would be explicitly reacting to the long-term spread, other than to final output and inflation objectives. The significance of the spread as a regressor in an estimated interest rate reaction function would then signal to what extent inflation and output stabilisation were sacrificed in the attempt to stabilise the exchange rate within the ERM band. In addition, the evolution of the way in which the central banks reacted to the spread and to other regressors, and an assessment of the stability of estimated reaction functions, would illustrate further aspects. For example, it would show the extent to which the adoption of a tougher exchange rate commitment since late eighties (the "hard ERM"), and the varying commitment of national authorities to programs of fiscal consolidation, affected short-term interest rate determination.

Estimated interest rate reaction functions for the countries in our sample show that budget policies had severe effects on, and critically constrained, the behaviour of monetary authorities. In all countries, monetary policy stances seem to have been often motivated by the need to respond to changes in the credibility of the country's exchange rate position within the ERM band. Interestingly, with the start of the "hard ERM" phase, in France, Belgium and Ireland, the importance of the long yield spread tends to decrease as severe efforts of fiscal retrenchment were undertaken. In such countries, the ERM turbulence in 1992-93 does not appear to have significantly affected interest rate policies, probably thanks to the largely achieved macroeconomic stability. On the contrary, for Italy, well-founded concerns surrounding its macroeconomic policies at the eve of Stage Three of EMU severely constrained interest rate determination. Some more consistent fiscal consolidation somehow eased the process of nominal convergence vis-à-vis the other Member countries, but only in latter part of the sample.

The chapter unfolds as follows. **Section 2** outlines the benchmark theoretical model used to derive the reaction function subsequently estimated. Clearly, despite the similarities between the interest rate equations estimated here and in the previous chapter, the presence of the exchange rate band make the flexible rates model previously adopted fully inadequate. In addition, one needs to take into consideration the alternative sources of exchange rate risk present during the ERM years and the way in which these affected interest rate policies. This is why **Section 3** illustrates some details of our estimation methodology, and shows how we tried to obtain a measure of exchange rate risks

[2] Or, alternatively of European median voters' attitude as regards the costs of disinflationary programs towards German

orthogonal to expected inflation. In **Section 4** we briefly examine monetary policy developments in the four countries in our sample, while in **Section 5** we finally present and comment our estimates. **Section 6** briefly summarises what we have found.

## 2. Modelling Interest Rate Rules in the ERM: A Simple Theoretical Framework

In this section, we modify the baseline theoretical framework studied in the previous chapter. Our purpose is to obtain an empirically testable relationship between the policy instrument and some final objectives. This will be placed in a policy framework within an imperfectly credible fixed-exchange rate system[3]. It would be trivial to simply include an exchange rate regressor in the baseline reaction functions estimated in the previous chapter. Indeed, this was done for some of the open economies considered in that case. In the present case, however, the existence of an exchange rate central par between the countries at hand makes such an option unfeasible, as interest rate policies under a fixed exchange rate regime do not respond to exchange rate movements in a linear fashion.

We refer to the second section of the previous chapter for a somewhat more extensive discussion of the basic forward-looking policy framework, which forms the basis for most recent monetary policy studies. There, we stressed the fact that a relatively broad consensus has emerged about a new wave of models, in which aggregate relationships are explicitly derived from the optimising behaviour of households and firms (Walsh, 1998; Rotemberg and Woodford, 1999; Bernanke, Gertler and Gilchrist, 1999).

The resulting behavioural relationships allow current aggregate values for macroeconomic variables to depend, *inter alia*, on the future course of monetary policies.

We postulate almost all the same aggregate relationships that we derived at the outset of last chapter's model. The main difference with what claimed there lies in the fact that we now explicitly introduce exchange rate considerations in the model and in the policy objective function. The way we carry out such task will enable us to model the effects of policy credibility on the current exchange rate, following recent developments

median voters' one.

[3] As we have mentioned in the Introduction, the numerous realignments negotiated during the ERM support the belief that the latter was, particularly during its early years, a quasi-flexible exchange rate system. See Giavazzi and Giovannini (1989), Fratianni and von Hagen (1992), Cukierman (1992), Garber and Svensson (1995), De Grauwe (1997), Dornbusch *et al.* (1998), Obstfeld (1998).

introduced by the well-known "fiscal theory of price level determination" (Canzoneri, Cumby and Diba, 1998; Cochrane, 2000).

Let $y_t$ and $y_t^*$ be again the current and potential level of output, and $p_t$ the price level. The following expectations-augmented Phillips curve is assumed (all variables but interest rates in logs)[4]:

$$\pi_t = \beta\pi^e_{t+1} + \varphi\left(y_t - y_t^*\right) \qquad [1]$$

In [1], current inflation depends on the inflation rate expected in the next period, and on the current output gap. In turn, the latter is affected by the deviations of the nominal interest rate from its expected value, and by a white noise shock:

$$y_t - y_t^* = -\lambda\left(R_t - R_t^e\right) + \varepsilon_t \qquad [2]$$

Again, [2] is a customary relationship whereby a positive surprise in the interest rate level negatively affects current output. The Fisher *ex ante* parity holds, so that

$$R_t^e = r^* + \pi^e_{t+1} \qquad [3]$$

Now, classical contributions on exchange rate bands (Miller and Weller, 1991; Flood, Rose, and Mathieson, 1991; Bertola and Caballero, 1992; Delgado and Dumas, 1993; Garber and Svensson, 1995) all stemming from Krugman's (1991) seminal work, implicitly allow for policy credibility effects in the behaviour of exchange rate[5]. The common starting point of such literature (see Miller and Weller's paper, for example) is a basic relationship in which the current exchange rate $e_t$ is determined according to the following (in logs):

$$e_t = m_t + v_t + \gamma_r \Delta_t^e e_t, \qquad [4]$$

[4] See Bernake, Gertler and Gilchrist (1999) or Rotemberg and Woodford (1999) for details on the derivation of all aggregate relationships.

[5] See also Bartolini and Prati (1999), and Avesani, Gallo and Salmon (1999)

where $m_t$ is money supply (or an endogenous policy variable), $v_t$ is the cumulative velocity of money, and $\Lambda_t^e e_t$ represents the instantaneous expected rate of change of the exchange rate. The latter affects the current value of the exchange rate through the semi-elasticity of money demand, $\gamma_r$. The last term on the right-hand side thus reflects market expectations about the future course of monetary policy, likely reflected in exchange rate risk considerations.

An ideal way of modelling the credibility effects stemming from ERM membership would call for incorporating [4] into an optimum model of policy behaviour[6]. Whereas such idea is certainly appealing and indicates an interesting direction for further theoretical research, our aim here is different. In particular, the use of [4] in an amended version of the model presented in the previous chapter would make the subsequent estimation of the derived policy rules hardly manageable.

More simply, we assume that the current value of the exchange rate is determined by the sum of two components: the differential between home country's and Germany's short-term interest rates, and a measure of the exchange rate risk. In what follows, we define the (log of the) exchange rate $e_t$ in terms of its deviations from the ERM parity with the German mark, which is in turn assumed to be, for simplicity, zero. The current exchange rate may then deviate from the central par according to[7]:

$$e_t = -s\left(R_t - F_t^G\right) + P_t \qquad [5]$$

In [5], we take the two interest rates used as the policy instruments in the home country and in Germany. We chose the Fibor rate as the foreign interest rate because it is strongly collinear, over the medium term, with the German call money rate, while its use avoids some likely simultaneity with home country's policy instrument.

Equation [5] is a rough-and-ready way of characterising the relationship between exchange rate expectations in each period and the credibility of the overall policy thrust. In particular, the equation is a linear approximation of a relationship that, given eq. [4], should be thought as non-linear in practice. According to it, the current exchange rate is affected by the current differential between domestic short interest rate and Germany's Fibor rate $F_t^G$, while it is driven away from the parity whenever the exchange rate risk $P_t$

[6] Coles and Philippopoulos (1997) represents a first and successful attempt.
[7] The superscript "G" denotes German variables.

is different from zero. While the rationale for such a law of motion is intuitive in our simplified context, it allows to model exchange rate dynamics avoiding exceedingly complex relationships. In particular, while it would be optimal to employ a relationship akin to [5], but closer to an uncovered interest parity, this would involve postulating some exogenously fixed mechanism of determination for exchange rate expectations. Here, our aim is different. The use of the above, *ad hoc*, relationship, will enable us to model the state of inflation and exchange rate expectations entirely within the model. Furthermore, its use allows us to characterise the influence of fiscal policy stance on future inflation expectations, without departing too much from the classical framework outlined by equation [4].

Finally, we assume that the exchange rate risk is fully reflected in the current long-term interest rate differential with Germany, $S_t = LR_t - LR_t^G$:

$$P_t = \gamma\left(LR_t - LR_t^G\right) \qquad [6]$$

Under normal circumstances, one may think that the long-term spread simply reflects expected differences in future inflation rates between the two countries. Favero, Giavazzi and Spaventa (1997) provide robust empirical evidence as to the determinants of the long yield spreads in Europe, and amongst these, exchange rate risk seems to explain well the recent historical behaviour of such spread against Germany.

As in the previous chapter, monetary policy authorities's objectives are modelled such as to involve stabilisation, in each period, of the deviations of current inflation and output from respective targets $\tilde{\pi}$ and $\tilde{y}$ (the latter, as in the previous chapter, is assumed to lie above the market-clearing level. See Barro and Gordon, 1983; Cukierman, 1992; Rogoff, 1985; Svensson, 1997a). Rotemberg and Woodford (1999) show that objectives like those sketched above can be derived directly from the minimisation of society's welfare function.

Moreover, we assume with Svensson (2000a) that the central bank attempts to minimise interest rate changes, and the departures of the policy instrument from its expected value. Finally, ERM membership entails that authorities' inflation target is given by German inflation[8]: $\tilde{\pi} = \pi_t^G$, and that central bank's loss function penalises deviations of the exchange rate from the zero parity:

$$L_t = \chi\left(\pi_t - \pi_t^G\right)^2 + \left(y_t - \bar{y}\right)^2 + \rho_1\left(R_t - E\{R_t\}\right)^2 + \rho_2\left(e_t\right)^2 \qquad [7]$$

Given such objectives, the central bank has a clear incentive to set its policy instrument below the level expected by the private sector[9], in the attempt to push output above its market-clearing level. However, the presence of the exchange rate band somehow shrinks such incentive, since, according to [5], a lower level of the short-term interest rate vis-à-vis the Fibor rate directly triggers depreciation. Participants in the foreign currency market, in turn, perceive the existence of this incentive, and adjust their expectations about inflation and the exchange rate accordingly.

Now, three scenarios are possible. In the first, called for simplicity Central Bank Dominance (CBD), monetary authorities rein in inflation and the exchange rate. Under such regime, the exchange rate parity is fully credible, i.e. the probability that the parity will be re-negotiated in the following period is zero. Under these circumstances, the exchange rate will not systematically diverge from the central par, because the expected inflation differential and the long-term interest rate differential between the two countries will not be different from zero. On the right-hand side of [5], the perceived exchange rate risk is zero, and the current exchange rate is unambiguously determined by the short-term interest differential. Since we assume that supply shocks are uncorrelated, there is no scope for systematic inflation differentials between the home country and Germany. The current exchange rate then behaves according to:

$$e_t^{CBD} = -s\left(R_t - F_t^G\right) \qquad [8]$$

Given that $E_t\left\{F_t^G\right\} = F_t^G$, the policy problem is to minimise [7] with respect to the policy instrument, subject to [1]-[3] and [8], and holding expectations as given. We obtain the following optimal feedback rule for the short-term interest rate:

[8] This is consistent with traditional models of the ERM. See Giavazzi and Pagano (1988), and Giavazzi and Giovannini (1989)

$$R_t^{CBD}=\left(\frac{1}{\Omega}\right)\left\{\pi_{t+1}^{e}\Gamma+r^{*}\left(\chi\varphi^2\lambda^2+\lambda^2\right)-\mathrm{I}\right\}+\frac{\left(\chi\varphi^2\lambda+\lambda\right)\varepsilon_t}{\Omega+\rho_1} \qquad [9]$$

where

$$\begin{aligned}\Omega&=\chi\varphi^2\lambda^2+\lambda^2+\rho_2 s^2\\ \Gamma&=\chi\beta\varphi\lambda+\chi\varphi^2\lambda^2+\lambda^2\\ \mathrm{I}&=\chi\varphi\lambda\pi_t^G-\lambda\left(y_t^*-\tilde{y}\right)+\rho_2 s^2 F_t^G\end{aligned} \qquad [10]$$

In equilibrium, $\pi_{t+1}^{e}=\pi_{t+1}^{eG}$, and the short-term expected interest rate equals the expected German Fibor.

In the alternative regime, called for convenience Fiscal Dominance (FD), the exchange rate band has zero credibility. The current exchange rate is not uniquely determined by monetary policy actions, because the bank is unable to control inflation. That is, under such a regime, there is a probability equal to one that a current fiscal shock will force the central bank to expand money supply in *t+1*, and that the parity will be re-negotiated[10]. In other words, the exchange rate risk is positive, and central bank's actions in *t* affect the level of the current exchange rate only to a marginal extent, overridden by the "unbalancing" behaviour of the fiscal authority.

Consequently, the exchange rate is expected to depreciate by an extent proportional to the exchange rate risk. We saw that such risk is reflected in the current spread between interest rates on home and German long-term bonds. The current exchange rate in [4] is thus determined as:

$$e_t^{FD}=-s\left(R_t-F_t^G\right)+\gamma\left(LR_t-LR_t^G\right) \qquad [11]$$

Under the present regime, the bank's optimal feedback rule becomes:

$$R_t^{FD}=\left(\frac{1}{\Omega}\right)\left\{\pi_{t+1}^{e}\Gamma+r^{*}\left(\chi\varphi^2\lambda^2+\lambda^2\right)-\mathrm{I}+\rho_2 s\gamma S_t\right\}+\frac{\left(\chi\varphi^2\lambda+\lambda\right)\varepsilon_t}{\Omega+\rho_1} \qquad [12]$$

[9] Which, in a perfectly credible equilibrium, would coincide with the Fibor.
[10] See Cukierman (1992), who, independently of the recent debate on the fiscal determination of the price level, illustrates some rationale for a monetary surprise in presence of a high level of public debt.

Note that [12] differs from [9] for the presence of the term in the long interest rate spread within curly brackets.

Finally, let us assume that, in a third scenario, in $t$ the market is uncertain about which regime in actually in place. In other terms, the exchange rate band is not fully credible. Market participants will assign some positive probability $q$ to the event that the current regime is, in fact, one of FD. In this case, the current exchange rate will be a weighted average of [8] and [11]:

$$\begin{aligned} e_t &= qe_t^{FD} + (1-q)e_t^{CBD} = \\ &= q\left[-s\left(R_t - F_t^G\right) + \gamma\left(LR_t - LR_t^G\right)\right] + (1-q)\left[-s\left(R_t - F_t^G\right)\right] \\ &= -s\left(R_t - F_t^G\right) + q\gamma S_t \end{aligned} \qquad [13]$$

By solving the optimal policy problem under uncertainty about the *fiscal* policymaker behaviour, we obtain a reaction function akin to [12], but this time with the parameter associated to the yield spread modified by the factor $q$:

$$R_t = \left(\frac{1}{\Omega}\right)\left\{\pi_{t+1}^e \mathrm{I} + r^*\left(\chi\varphi^2\lambda^2 + \lambda^2\right) - \mathrm{I} + q\rho_2 s\gamma S_t\right\} + \frac{\left(\chi\varphi^2\lambda + \lambda\right)\varepsilon_t}{\Omega + \rho_1} \qquad [14]$$

Under such regime, expected inflation is a weighted average of expected inflation prevailing in the FD and CBD scenarios. Obviously, to the extent to which the expected inflation differential is reflected in the long-term spread, the latter should be collinear with expected inflation. We show below how we tried to avoid this collinearity problem when estimating our reaction functions. In a multi-period framework, one can assume $q$ as revised in each period according to the observed central bank's behaviour, that is, agents revise $q$ according to past exchange rate and inflation volatility. Following this, one might try to estimate some version of eq. [14].

In line with the analogous assumptions of the previous chapter, in the estimation of [14] we allow for imperfect information as to central bank's objectives[11], and we also assume that the latter's ability to predict the supply shock is limited. Following Muscatelli

[11] Also assumed to be time-varying.

(1998, 1999), Faust and Svensson (2000), and Walsh (1998), we let private sector's beliefs about central banker's relative inflation aversion to be represented by:

$$\chi_t^e = \phi\chi_{t-1} + \kappa_t \qquad \kappa_t \sim (0, \sigma_\kappa^2) \tag{15}$$

In addition, we suppose that the policymaker has only limited knowledge of the state of the economy (the supply shock $\varepsilon_t$), that he/she makes inferences on it through a forecasting process, and that such forecast, $\varepsilon_t^f$, is private information. In each period, then, private agents are uncertain as to whether the shock they observe is due to a true supply shock, or it simply reflects a shift in policymaker's preferences $\kappa_t$ (Cukierman, 1992). Private sector's perception of the interest rate rule will thus be different from [14], and will be:

$$R_t = c_0 + c_1 F_t^G + c_2 \pi_{t+1}^e + c_3 \varepsilon^f + c_4 S_t + \xi_t \tag{16}$$

In [16], the parameters are linear functions of those in equation [14], but now $\chi_t^e$ has replaced $\chi_t$, and the supply shock is only the forecast of the one on the right-hand side of [14]. Note also that the coefficient $c_4$ is a function of $q$, the exogenous probability assigned by the market to the likelihood of a regime of FD. Agents update their expectations about the business cycle and central banker's preferences in each period, by looking at past disturbances' variances.

In the case of a policy break, like the move to narrower exchange rate bands, the announcement of a tougher commitment to the parity, or the adoption of some fiscal convergence criteria, like those contained in the Maastricht Treaty, we have two possible scenarios. If the new regime is a fully credible one, the adjustment of equilibrium inflation and interest rates is immediate. Moreover, if the fiscal policy stance does not suggest a future switch to fiscal dominance, the coefficient $c_4$ in [16] should be close to zero. If the reform is instead only partly credible, nominal variables will adjust gradually to the new steady state. Assuming one can obtain estimates for the inflation and output regressors of the above relationship, significant and permanent changes of estimated coefficients in [16] could be easily detected. These, along with eventual instability of the overall equation in correspondence of major policy shifts, would signal either changes in policymakers' preferences, or the introduction of some institutional reforms, or both[12]. In addition, the size and the behaviour of the estimated coefficient $c_4$ will signal whether interest rate policy has in fact targeted the exchange rate risk as perceived by market participants. In other words, with a perfectly credible central par, the central bank can exploit the exchange rate band to pursue domestic objectives. With less than full credibility, the current position of the exchange rate within the band signals the overall credibility of the exchange rate commitment. In the latter case, the central bank is ready to offset the effect of a positive exchange rate risk on the current exchange rate through interest rate changes not otherwise triggered by changes in the level of economic activity.

To sum up, what we aim to obtain, by studying estimated versions of reaction functions like [16], is an assessment of central banks' conduct in France, Italy, Belgium and Ireland on the road to their participation to EMU. Such conduct can be usefully exemplified by estimated rules like [16], which capture the implicit way in which monetary policies translate in a simple rule expressed in terms of expected inflation, output gap and the long-term spread (in the spirit of Batini and Haldane, 1999).

As in the preceding chapter, we are not dogmatic about the functional form to be estimated. We shall include additional regressors to the baseline specification in [16]. In all the countries we examine, we will test whether variables like lagged money growth, or the change in central banks' official reserves excluding gold, significantly enter the policy rule[13]. Clearly, what matters when one evaluates the effectiveness and stability of policy rules, is their performance in terms of announced final, not intermediate, objectives. Should any of the above indicators enter an estimated reaction function, one would then conclude that the role it plays in the policy rule is similar to those played by final inflation, exchange rate and output objectives. Again, should we really find a significant role for money growth in the estimated interest rate function of any of the countries we study[14], we would conclude that the stabilisation of money supply in that country took place at the expense of meeting final output, inflation, and exchange rate objectives.

The key difficulty with the estimation of reaction functions like [16] is represented by the availability of the unobserved series for expected inflation and potential output (Clarida *et al.*, 1998; Gerlach and Smets, 1999; Favero and Rovelli, 1999), along

[12] Of course, a third possibility would be that the observed instability is simply due to changes in the underlying behavioural relationships between the variables of interest. The way in which we calculate the series for expected inflation and the output gap, however, takes into account such possibility.
[13] In the case of reserves, changes in their level significantly enter all estimated interest rate reaction functions.
[14] This was the case with Italy over the latter part of the sample, as we show below.

with the identification of some updating mechanisms for all expected variables. In the previous chapter we showed that one optimal, though not unique, way of solving this issue, is applying the Kalman filter to such variables. As we adopt this technique here again, we refer to the brief discussion of such choice, and of the recent literature on monetary policy rules, contained in the previous chapter.

## 3. Econometric Methodology

### *3.1 Measuring Inflation Expectations and the Output Gap*

Given the uncertainty with which the central bank is assumed to observe the behaviour of output and inflation, we adopted an unobserved component approach to the derivation of trend inflation and output[15]. In particular, the calculation of our measures of expected inflation and potential output involved fitting a Structural Time Series model (STS, Harvey, 1989) to a univariate[16] specification of inflation and real GDP processes[17]. The STS methodology allows to find an appropriate decomposition of the original series into (stochastic) trend and cyclical components, using a Kalman filter estimation procedure. The latter constitutes a natural way of allowing model parameters to be optimally and gradually updated, as new information becomes available. This way, the model better reflects the gradual learning by private agents' -and central bank's- of the behaviour of inflation and output processes, as new information becomes available. Furthermore, it allows incorporating private agents' learning about central bankers' preferences into the estimation strategy we pursue. As we have seen, this aspect is critical in a context of imperfect information like ours. ML estimation of the hyperparameters of interest then ensues. In the present case, both series contained marked cyclical components, which were modelled by allowing for one or two stochastic cycles, as needed[18].

[15] For a similar, though not identical, application of the unobserved component approach to the estimation of the output gap in the EMU area, see Gerlach and Smets (1999).

[16] As in the previous chapter, we tried a multivariate specification of the model. The series were allowed to respond to simultaneous and lagged innovations of other macroeconomic variables, but the explanatory power of such specifications was never greater than in the univariate case. We then definitely chose the latter. Further details on the generation of regressors and the econometric methodology are available from the author upon request.

[17] Again, in Chapter 1 we extensively motivate this choice, and discuss some possible alternatives.

[18] The STAMP 5.0 software was used to estimate the STS models, through the conventional concentrated diffuse likelihood technique. Output and inflation were found to be I(1), and to have significant cyclical components. For a similar approach to forecasting inflation in the presence of potential structural breaks, see Stock and Watson (1999). The trends were specified according to a "stochastic level, damped slope" formulation, which however did not yield substantial differences relative to the "fixed level, stochastic slope" case.

We estimated quarterly models for real GDP and inflation for each country, obtaining a decomposition of the series into trend, cycle, and irregular components. In the case of GDP, a convenient decomposition was generated by applying the Kalman filter on the trend component. The latter was then optimally computed based on one-step-ahead predictions of the state vector. This way, estimates of potential output are based only on past information, rather than on the full sample.

In the case of inflation, we simply computed one-step-ahead prediction errors from a univariate STS model to obtain a measure of expected and unanticipated inflation. This way, model parameters are updated only gradually, as new information becomes available. The use of the basic filter, as opposed to the smoothing algorithm (Kim and Nelson, 1999), guarantees that future observations never affect the calculation of the stochastic components.

We finally turned to the estimation of the interest rate reaction function [16], for each country, using simple Recursive Least Squares[19]. Recursive estimates are particularly useful when, as in the present case, one needs to observe how estimated coefficients evolve over time. Moreover, we performed conventional structural-stability tests on the residuals of each equation, to capture eventual signs of breaks.

We estimated the quarterly models for the four economies over a sample starting in 1980Q1 and ending in 1997Q2. The data we use are quarterly observations taken from OECD Main Economic Indicators and IMF International Financial Statistics, as described in the Data Appendix.

### *3.2 Interest Rate Differentials and Expected Inflation*

As explained in Section 2 and in the Introduction, one of our aims is to understand the effects that the credibility of the fiscal stance had on interest rate policies. This task can be tackled in many fashions. An ideal one would be assessing the alternative effects of the business cycle and fiscal policies on the monetary stance. Some recent work on the effects of the fiscal stance argues that the latter may pose additional constraints on the way in which inflation is generated (Canzoneri, Cumby and Diba, 1998). In particular, such literature draws a distinction between regimes of central bank dominance (CBD) and

[19] Estimated coefficients are computed using the author's GAUSS code. Stability tests are conducted using PcGive 9.1.

those of fiscal dominance (FD). In the latter, primary surpluses are not responsive to the level of public debt, so that the price level and the money stock need to adjust to ensure fiscal solvency[20]. Inflation simply adjusts to the needs of fiscal solvency. Under a CBD regime, instead, primary surpluses systematically react to the level of public debt, and inflation is determined according to central bank's unconstrained optimal feedback rule for money supply and interest rates. Unfortunately, we are not aware of any significant application of this interesting literature to an open-economy context.

Melitz (1997) provides some empirical support to the idea that, over the sample we are studying, the complementarities between budget and monetary policies within European countries (and elsewhere) were substantial. In other words, there appears to be trace of *tout-court* fiscal dominance in none of the former ERM countries. In principle, however, the strategic interactions between fiscal and monetary policies -and authorities- in presence of budget rules as those endorsed by the Maastricht Treaty, are quite complex. Very recent work (Leith and Wren-Lewis, 2000) has shed some light on the issue of whether the above-mentioned "fiscal theory of price level determination" (Woodford, 1995; Canzoneri and Diba, 1996; Canzoneri, Cumby and Diba, 1998; Cochrane, 2000) helps explaining some of these theoretical features. Some more sophisticated empirical research is however needed, and with reference to individual ERM countries, as well as for the EMU area as a whole, under way.

In the previous section, we illustrated a simple way in which the interactions between the public and the monetary authorities in presence of an exchange rate agreement may unfold. Our theoretical model says that the spread between the interest rate on long-term bonds relative to Germany can be employed as a measure of the overall credibility of economic policy stance, as perceived by financial markets participants. More precisely, we assumed that the long-term spread might reflect the market view about the likelihood that national authorities may be forced to expand monetary policy, thus loosening control of inflation and the exchange rate. The larger such probability, the larger the spread. If the spread systematically enters an estimated version of our final reaction function in equation [16], we then conclude that interest rate policies in that country have been reacting to the perceived exchange rate risk associated to the country position within the ERM band. We can then view such simple mechanism as one that guarantees an observable link between the behaviour of fiscal authorities and that of monetary

[20] That is, to ensure that government's intertemporal budget constraint is satisfied.

authorities. The size and the behaviour over time of the coefficient attached to the spread in the estimated reaction function then describe the extent to which the fiscal stance constituted an additional constraint on national interest rate policies.

However, there are intuitive reasons to expect that long-term interest rates differentials and expected inflation are strongly collinear. Using the pure interest rate differential as a regressor in our estimates would involve dealing with an aggregate measure of the exchange rate risk that is clearly correlated with expected inflation.

Favero, Giavazzi and Spaventa (1997) study the daily behaviour of the spread on the 10-year benchmark bonds of Italy, Spain, Sweden, and Germany[21]. They identify and measure three components of such spread: one directly related to the expectation of exchange rate depreciation, another due to differences in tax regimes across countries, and one which reflects the market assessment of default risk. While it would be optimal to consider such complex decomposition in our work, its use would invariably complicate the subsequent estimation of reaction functions. Therefore, we limit ourselves to the calculation of a measure of the exchange risk more closely associated with the first and third component identified by Favero *et al.* (1997). We attempt to do so by purging the component of the total differential directly associated with expected inflation, and obtaining a proxy for the part of exchange rate risk orthogonal to inflation expectations. Simple recursive regressions of the interest differential on expected inflation, for each country, appeared to serve the purpose without significant bearings on the precision of our final estimates. Residuals from those recursive regressions (*Adjspread*) were then inserted as regressors into our baseline reaction function, along with expected inflation, the output gap, and other regressors.

Though difficult to interpret, the results of these preliminary regressions (**Table 1**) are as expected. The coefficients are all strongly significant and correctly signed.

Finally, as we have already mentioned, we explicitly account for the possibility that the central bank might have responded to changes in intermediate objectives not included in our baseline specification. The inclusion of additional regressors like money growth and the change in central bank's foreign exchange reserves excluding gold, is tested for each equation. Those additional variables would not be relevant simply because they provide information about future inflation and real activity. This will instead allow us to check whether such variables had some relevance in the formulation of interest rate

[21] We calculate the spread on the same category of bonds.

policies, beyond the effects that they might have had on the generation of inflation expectations, the exchange rate risk, and the output gap.

To better fix ideas before commenting on the results of estimated interest rate reaction functions, we briefly examine some of the major developments involving the economic policies of each country during the period covered by our estimates.

## 4. Monetary and exchange rate developments in France, Italy, Belgium and Ireland

According to many historical accounts[22], the ERM was not initially established as a rigid system of fixed exchange rates. The universal perception was that each country's parity could have been adjusted according to accommodate changes in underlying economic conditions. In fact, during the first four years, seven (out of twelve) realignments took place. As time went on, however, the system evolved towards a more rigid regime, and the years between 1987 and 1992-93's breakdown (the so-called "hard ERM" phase) witnessed no adjustment[23] at all. Since 2 August 1993, the bilateral margins around the exchange rate parities were widened from ±6 to ± 15%.

What follows is a concise history of the interplay between domestic macroeconomic conditions and the external constraint represented by the ERM, in the group of countries under study.

The Banque de France has repeatedly argued that since late '70s its policy had relied on two fundamental intermediate objectives: strict adherence to the ERM, and money supply growth (Fratianni and Salvatore, 1993; OECD, 1999c). Since 1977, the Bank has set targets for monetary growth: M2, from 1988 to 1990, M3 thereafter.

During the first years of French participation to the EMS, the commitment to the exchange rate target seemed a relatively loose one, and capital controls were heavily used to shield domestic money markets from "undesirable" fluctuations. In 1984, however, the overall policy thrust turned unambiguously more anti-inflationary, and the activation of a stricter targeting of the exchange rate, rather than of money growth, gradually took place. The "hard ERM" phase in early '90s saw France's attempts --as well as other countries' ones-- to seek further convergence of domestic inflation levels with Germany.

[22] See Fratianni and von Hagen (1992), amongst others.

Consequently, between 1984 and 1990 capital controls were progressively dismantled. However, the run-up to closer monetary co-ordination within the EMS broke down during 1992-1993, when the franc was forced outside "strict" targeting of the DM. Despite that, the Banque de France managed to control inflation and the exchange rate. According to an interesting view of the events following the ERM break-up in 1992-93 (Bartolini and Prati, 1999), those events changed the exchange rate strategy of the Bank. Its policy of tolerating short-lived fluctuations of the DM/FF rate, while still strongly committed to longer-term exchange rate parity, narrowed the scope for short-run speculation[24]. Since the overall stance of economic policies appeared consistent with the pursuit of a rigid exchange rate in the long run, that strategy helped in stabilising inflationary expectations as well.

From an operational point of view, the mid-eighties saw a major change in the conduct of French monetary policy (Melitz, 1993; Mojon, 1998). A number of financial reforms progressively abolished the regime of administrative credit rationing and the day-to-day official fixing of the interbank rate prevailed up to 1987. Moreover, the Banque de France began to provide liquidity to the system acting essentially on the interbank and overnight money markets. Until the start of EMU's Stage Three, repurchase agreements were the main source of central bank money.

Finally, it is important to note that the Bank was granted full legal independence in 1993.

Since the split of 1981 with the Treasury, the Bank of Italy gained a substantial amount of formal independence. This in turn enabled it to gradually switch, as in the French case, from the use of credit ceilings to standard interest rate policy[25]. In recent years, the interest rate on repurchase agreements seemed to have become the main policy instrument (Gaiotti, 1999). In 1984, the Bank announced the first M2 official target[26]. It soon became clear that, however, monetary targets were subordinate to the government-set exchange–rate target, and that eventual harmonisation between the two had still to pass through restrictions to capital flows.

[23] Aside from the narrowing of Italian lira's band, in 1990.

[24] Anthony and MacDonald (1999) find some empirical evidence supporting this view. Their work shows that the mean-reverting properties of various ERM exchange rates were essentially the same with the broad band as with the narrow band.

[25] See Spinelli and Tirelli (1993), Fratianni and Spinelli (1997), Gaiotti, Gavosto and Grande (1997).

Italy had joined the ERM as a founding member in 1979, though the initial wider band, several realignments and strong capital controls made the exchange rate constraint relatively flexible throughout the first half of the eighties. Clearly, full financial liberalisation occurred only when (early 1990s) a more favourable climate of confidence about inflation control temporarily relieved the pressure on the exchange rate. At the same time, the lira entered the narrow band of the ERM. A loose fiscal policy stance and the mounting public debt, however, have cast a recurrent shadow on the ability of Italian monetary authorities to control nominal variables[27]. Thus, concerns became to grow over the compatibility of the current state of public finances with EMU provisions (the Maastricht Treaty was signed on 7 February 1992), if not with medium- to long-run overall sustainability[28]. The dramatic exit of the lira from the ERM in 1992 appeared, *inter alia*[29], as a direct consequence of such concerns.

Meanwhile, successful agreements on labour costs in 1992-93 had contributed to ease the pressure on inflationary expectations. However, the flight to foreign currency-denominated assets that accompanied the currency crisis was halted only when decisive steps towards a badly needed fiscal correction were finally undertaken by the second half of the '90s. By then, a more optimistic outlook for public finances probably contributed to somehow lower the risk premium on lira-denominated assets. In November 1996, Italy rejoined the ERM, and in 1998 the Bank of Italy became one the 11 founding Members of the ECB.

Belgian monetary authorities have always argued that in a small open economy the relationship between the exchange rate and inflation was far more stable, and reliable, for policy purposes, than the growth of monetary aggregates (targeted by Germany's authorities since 1974). Consequently, since the collapse of Bretton Woods, Belgium (along with the Netherlands), had joined various exchange rate arrangements in the

[26] The official intermediate objective of the Bank had previously been total domestic credit. This, as discussed in Spinelli and Tirelli (1993), and Fratianni and Spinelli (1997), entailed large crowding-out of private-sector credit and lack of control on monetary aggregates, in presence of large government budget deficits.
[27] For an effective assessment of the effects of these considerations on currency markets, see Giorgianni (1997).
[28] The ratio of central government deficit to nominal national income almost doubled in the decade 1981-91 relative to the previous one (Fratianni and Spinelli, 1997).
[29] One of the other most commonly argued causes unfolds as follows. The Italian participation to the ERM witnessed the accumulation of substantial inflation differentials between Italy and the other participating countries. Occasional parity realignments never fully compensated for such differentials. This way, the lira appeared, in the early '90s, as considerably appreciated, in real terms, vis-à-vis the other ERM members. The 1992 breakdown was then produced by long-term competitiveness difficulties, and directly triggered by the interest rate shocks following the German reunification. This argument, however, does not take account the existence of diverging productivity trends between Germany and its ERM partners (for such considerations, see, for example, Canzoneri, Cumby, Diba and Eudey, 1998)

attempt to provide a nominal anchor to its economy[30]. Aside from a devaluation of 8.5% in February 1982, in the 1980s monetary policy was essentially designed to maintain stable exchange rates between the franc and the ECU. In 1990, monetary authorities finally declared their intention to peg the currency against the D-mark within a very narrow range of fluctuation.

As in Italy, substantial budget imbalances generated relatively high real interest rates throughout the Belgian participation to the ERM. However, a decisive strategy of fiscal consolidation, pursued over the medium term, has progressively boosted confidence in the currency and overall policy credibility, then narrowing the scope for speculative attacks (see IMF, 1998; Perotti, Strauch and von Hagen, 1998). Moreover, the stability-oriented monetary policy of the seventies and eighties managed to curb inflation towards German levels already since mid-eighties. That strongly contributed to the decline of interest rate differentials with Germany, and to a steadily credible climate surrounding economic policies.

Until 1979, the prospects of the Irish monetary policy were closely tied with those of Britain, since Ireland had adopted a currency board fixed on the British sterling. This resulted, amongst other things, in a significant depreciation of the Irish punt against many "snake's" currencies. The entry of Ireland in the EMS in 1979, however, did not result in immediate convergence of domestic inflation on German levels. The strong trade links with Britain meant that the domestic price level was still closely tied to the British one, and that the persistent real appreciation of sterling was affecting Irish competitiveness as well. Moreover, as in the case of Italy and Belgium, substantial budget imbalances developed over the years have recurrently put the currency at risk of speculative attacks. However, the severe macroeconomic adjustment carried out since 1984 did start to produce some effects in the second part of the decade. In fact, between 1987 and 1989 the differential with German long-term interest rates dropped, as the main consequence of a more optimistic economic outlook and increased credibility of the fiscal stance on asset markets. Subsequently, a mix of tax cuts, parity realignments, and wage moderation boosted competitiveness, which stimulated a further economic expansion since 1994. This particular chain of events is now commonly classified as one of the few cases of "expansionary fiscal adjustment" depicted in the recent literature on budget consolidation

[30] See National Bank of Belgium (1998) for a summary of the history of Belgian franc.

(see Alesina and Perotti, 1997; Giavazzi, Jappelli and Pagano, 2000)[31]. The monetary authorities then left the punt to significantly appreciate vis-à-vis the D-mark, strategy subsequently reversed in 1997, and in 1998, when a 3% revaluation of the punt was the last official realignment of ERM history.

## 5. ERM in Action: Monetary Policy Reaction Functions

In Section 2 we saw that a simple policy reaction function can be derived within a standard sticky price model with fixed –but "adjustable"– exchange rates. Moreover, the class of policy rules in equation [16] is broadly consistent with the inflation-forecast based rules recently advocated in the monetary policy design literature (Svensson, 1997b, 2000a; Rudebusch and Svensson, 1999; Batini and Haldane, 1999). Here we simply refer to the considerations contained in the previous chapter for a discussion of such rules.

Bearing in mind the events summarised in the previous section, we then turn to illustrating our estimates of the interest rate reaction functions for France, Italy, Belgium and Ireland, using quarterly data for each country. The policy instrument we adopted for each central bank has been chosen following widespread opinions in the literature on the transmission of monetary policy impulses, and in all cases but Italy coincides with the call money rate. Further details on the single series are contained in the **Data Appendix**. The sample chosen is 1980Q1-1997Q2 for all countries.

In principle, congruent models of monetary policies during the EMS should be able to unveil some systematic link between national interest rate policies and inflation and interest rates in Germany[32]. Our results confirm that during the "hard ERM" phase of the EMS the leeway for the countries under investigation to carry out independent monetary policies, had overall shrunk, as one would expect. Policies do appear to have converged over time towards the behaviour of interest rates in Germany, and to have taken the exchange rate commitment with a progressively tougher attitude. Such shift, however, has been relatively gradual. Some country can be seen as adapting its national policies in order to achieve a swift, though not painless, convergence towards low and stable inflation. Others have probably suffered from the presence of structural hurdles that

[31] For a short but effective account of those Irish events, see Obstfeld (1998).

initially prevented full monetary co-ordination with the rest of Europe. Our measure of credibility, i.e., the adjusted spread on long-term bonds vis-à-vis Germany, significantly enters all estimated reaction functions, whereas the output gap never does so. On the other hand, the generalised significance of the coefficient attached to changes in foreign exchange reserves provides us with some additional evidence in favour of exchange rate considerations playing a substantial role in each central bank's policy. This overall confirms that the ERM bands were not a fully credible exchange rate regime, and that fiscal imbalances may have played a significant role in interest rate determination.

The illustration of our findings for each country proceeds as follows. **Tables 2** and **3** displays estimates for the long-run solved static reaction functions, while the recursive graphs in **Figures 1-4** show estimated coefficients –between 2-standard error bands[33]– and Chow's tests of structural stability. The study of single recursive coefficients' path over time can provide for an educated guess of possible shifts in monetary authorities' preferences. However, the relevance of such changes can be fully gauged only with reference to their impact on the estimated reaction function as a whole.

Our estimates for **France (Table 2, Figure 1)** show that monetary authorities have been shadowing German policies at least over the last decade. The coefficient associated with the 3-month German Fibor rate is significant and larger than the one on domestic expected inflation. Clearly, French authorities appear to have signalled their commitment to an anti-inflationary stance through a close shadowing of German interest rates. In parallel, domestic inflation considerations have apparently played a minor role in the setting of French interest rates in recent yeas. Recursive graphs show how this pattern has gradually but firmly been enhanced: after 1985, the coefficient on domestic inflation steadily falls, while the one on the German rate rises by nearly the same extent. The (adjusted) spread on German bunds seems to play an important role during the "soft ERM" phase. The gradual relative reduction of French fiscal activism, however, may have contributed to the decline of this coefficient, as the graph clearly shows. Interestingly, the coefficient associated with the changes in reserves, tough small, is always significant, whereas the growth of any monetary aggregates does not significantly enter the equation.

Overall, the estimated equation shows some sign of instability in correspondence of the 1992-93 ERM crisis. Since then the coefficients have slightly larger standard errors,

[33] An attempt in this sense, involving EMU-wide measures of output gaps, can be found in Gerlach and Smets (1999). Gerlach and Schnabel (1999) instead perform a brief exercise aimed at estimating a Taylor rule for the EMU area centred on the latter part of our sample.

and the 1up Chow test explicitly displays a break around those events. We interpret this as evidence that the uncertainties surrounding the start of Stage Three of EMU had some bearing on French interest rates' behaviour. Nevertheless, the size of these fluctuations is relatively small, compared to other countries in the sample -notably Italy- and the N-down Chow test is unable to pick up significant disturbances.

After a cursory look at these results, it is plain to say that the whole thrust of French monetary policy has turned progressively more inflation-averse since mid-eighties. The deceleration in inflation that ensued turns out to be the likely consequence of a change in central bank's and/or public's attitude towards inflation, as well as of some underlying structural change. As we saw in the previous section, structural adjustment replaced output stabilisation as a policy priority, while policies aimed at the liberalisation of prices and the lifting of capital restrictions were enforced in parallel with a less expansionary budget stance. A closer pegging of the franc vis-à-vis the D-mark made French interest rates progressively more sensitive to monetary developments in Germany. Such results also show that the "soft" exchange rate bands implemented since the 1992-93 crisis did not entail significant loosening in inflation control (as observed by Bartolini and Prati, 1999; see also Anthony and MacDonald, 1999).

Estimates of the policy rule for **Italy** (**Table 2, Figure 2**) find a significant coefficient on domestic expected inflation, the highest value in our country sample. In addition, the coefficient on Germany's Fibor, contrary to all other countries, overall is barely significant, and its size, as displayed in the recursive graphs, tends to shrink substantially after 1989. Some rationale to these results can perhaps be found by looking at the relevance of exchange rate risk considerations: the coefficient on the spread against German bonds is strongly significant and outweighs all other parameters. Recursive graphs show that its size seems to have grown in parallel with the well-known concerns about Italy's budgetary position. This points towards fiscal policy having peculiarly heavy effects on market perception about monetary authorities' ability to control inflation. According to the theoretical scheme laid out in Section 2, if the central bank is able to control inflation and exchange rate fluctuations in the long run, the current exchange rate depends on the expected sequence of short-term rates and exchange rate risk considerations. These are in turn affected by expectations about the business cycle, and by the belief that the Central Bank will intervene to stabilise the exchange rate. On the

[33] Recursive estimates are obtained with a GAUSS code, and plotted using GiveWin. Stability tests are from PcGive 9.1.

contrary, if fiscal authorities do not stabilise public debt, the current exchange rate becomes independent from the current actions of the Bank, and it is instead affected by the belief that monetary policy will sooner or later be forced to create a regime of high inflation and realign the parity. In this case, the long-term spread measures fiscal policy credibility, and the size of the coefficient attached to it in our estimated policy rules reflects the likelihood that markets attribute to such event. We believe that our findings for Italy strongly support the latter scenario.

In all instances, 1992-93 appears as a turning point for monetary policy. Around early 1990s, the record levels of public debt of the past decade had clearly undermined the overall credibility of economic policies. Furthermore, the resilience of Italian fiscal policymakers in pursuing a policy of "benign neglect" towards the state of public finances between late eighties and early nineties likely reinforced this tendency (Fratianni and Spinelli, 1997). Consequently, the spread of Italian long-term bonds vis-à-vis their German analogues simply inflated. In the presence of restrictions on capital flows and wide exchange rate bands, this might still be consistent with some form of exchange rate co-ordination. But once such "allowances" were lifted, the only chance an independent Banca d'Italia had of controlling nominal variables was to restrict interest rate policy to offset the impact of unbalanced fiscal policies on exchange-rate risk. This is why our adjusted measure of the interest rate differential, a direct indicator of fiscal policy credibility, turns out to be even more relevant than domestic inflation for interest rate setting.

In line with this interpretation, our estimates clearly signal a structural break around the 1992-93 ERM crisis. Since those turbulent events, the overall thrust of economic policies appears more firmly oriented towards achieving price stability and budgetary consolidation (OECD, 1999b).

Interestingly, Italy is also the only country in the sample in which the growth of a monetary aggregate (M1 in this case), though only since recent years, significantly enters the central bank's policy rule. Prior tests did not reveal M2 growth, the announced intermediate target, as a significant regressor. M1 is probably the most relevant monetary aggregate if pure transactions are targeted as *intermediate* objective of monetary policy. Nonetheless, the majority of existing estimates of policy reaction functions agree on the empirical irrelevance of such a case[34]. In addition, it is clear that in periods of rapid inflation fall, M1 growth is strongly affected by adjustments in real money holdings, and it becomes completely unreliable as an indicator of economic activity[35].

Turning to the two smaller countries in our sample, estimates for **Belgium (Table 3, Figure 3)** appear broadly in line with those obtained for France. The coefficient on expected inflation, however, is small and barely significant. At the same time, while the evident significance of Germany's Fibor highlights the announced strategy of convergence on Bundesbank's interest rate policy, the adjusted spread against German bonds plays a prominent role in central bank's reaction function. This is particularly evident in the early part of the sample, as the coefficient significantly shrank since late 1980s onwards. Interestingly, the rise in the size of Fibor's estimated coefficient is almost perfectly mirrored by the fall in the coefficient on expected inflation and the adjusted spread. The short-term interest rate seemed to react, in line with results on all other countries, to the change in reserves as well. A glimpse at the recursive graphs shows that the country's interest rate reaction function was quite stable over time: very narrow confidence bands around all coefficients, little evidence of significant breaks, even around the 1992-93 crisis. That is, our estimated policy rule shows an overall stable relationship between interest rates, inflation, and the exchange rate. Belgian monetary policy could have suffered, as the Italian one, from the presence of a record level of public debt, and the subsequent rise in exchange rate risk premia. However, nominal convergence with Germany seems to have been pursued more resolutely than in the Italian case, and fiscal consolidation was successfully achieved over a medium- to long-term horizon (OECD, 1999). Both processes substantially staved off speculative attacks, and further enhanced exchange rate stability.

The picture sketched by the estimated reaction function for **Ireland (Table 3, Figure 4)** is somehow halfway through what we found for Belgium and Italy. The coefficient on expected inflation highlights the importance of domestic inflation

[34] The most orthodox example of monetary targeting regime, Germany, is now seen in the literature as an implicit inflation targeter (inter alia, Bernanke, Laubach, Mishkin, and Posen, 1999). The majority of the available evidence points out that monetary aggregates are never significant in estimated reaction functions for the G3 and the inflation targeting regimes (Clarida, Gali and Gertler, 1998, and the previous chapter). In any instance, however, broad rather than narrow monetary aggregates are universally indicated as the ones with more desirable stability and controllability properties (see Friedman and Kuttner, 1996, or Friedman, 1996, for a detailed account of the failures of monetary targeting experiments in the US).

[35] See the extensive discussion in Chapter 3. With reference to the present case, an advocate of monetary targeting in Europe argues: "When a country attempts to maintain a fixed value of the currency vis-à-vis the DM, internal monetary developments become theoretically highly endogenous, and almost unmanageable for the monetary authorities" (Groeneveld, 1998).

stabilisation as a final objective of monetary policy, whereas the output gap, in line with what we expect for a country that experienced sustained fiscal expansions, is never significant.

Changes in official reserves play a significantly stronger role than in other countries. This points to their role as to a short-run absorber for shocks to the exchange rate, given the recurrent misalignments triggered, *inter alia*, by the exceptional economic growth of the last decade.

Peculiarly, both coefficients on the long-term spread and the German Fibor do not significantly differ from one in recent years. This signals a very strong attitude on the part of the Central Bank of Ireland towards shadowing Bundesbank's interest rate policies, while trying to offset the ups and downs of a prolonged fiscal consolidation (see Alesina and Perotti, 1997; Perotti, Strauch and von Hagen, 1998). The coefficient on the adjusted spread was very large until 1989-1990; thereafter the severe fiscal retrenchment enacted -- and facilitated by strong real growth- likely made it converge towards lower levels.

Overall, the policy rule appears relatively stable in recent years: confidence bands are relatively narrow around estimated coefficients, while the only peak displayed by one of the structural stability tests is found in correspondence of the 1992 currency crisis. Since then, the behaviour of interest rates in Ireland became even more synchronised with ERM's core, although the sustained economic growth, as we saw, required some additional exchange rate adjustment.

## 6. Concluding Remarks

In this chapter, we have examined the behaviour of interest rate policies in France, Italy, Belgium, and Ireland, during the eighties and on their more recent road to EMU. Estimated reaction functions for each country provided us with some evidence about the relative costs these countries faced in maintaining the parities within the ERM, and in fulfilling the basic convergence criteria laid out in the Maastricht Treaty. The role of exchange rate risk, and the way long-term interest rate differentials reflected the overall credibility of fiscal policies in each country, has been analysed in relation to the trade-offs national policymakers faced between domestic and external objectives. By allowing each country's policy rule to reflect the extent to which exchange rate risk affected interest rate determination independently from expected inflation, we obtained an indirect measure of how fiscal policy credibility influenced central bank's decisions.

Our findings highlight quite clearly that in some cases (France, Belgium) the process of convergence towards a stable exchange rate vis-à-vis Germany and a tight inflation control took place amidst sustained policy reforms over a medium- to long-term horizon. In addition, the ERM turbulence in 1992-93 does not appear to have impaired the largely achieved macroeconomic stability. On the contrary, Italy's, and to a much lesser extent, Ireland's monetary policies, seem to have been severely constrained, in their efforts to stabilise the economy, by the lack of credibility of their respective fiscal stances. For Ireland, such pattern appears to have been promptly reversed, likely with the help of a very favourable growth outlook. For Italy, well-founded concerns surrounding its macroeconomic policies at the eve of Stage Three of EMU severely affected interest rate policies. Some more consistent fiscal consolidation somehow eased the process of nominal convergence vis-à-vis the other Member countries, but only in latter part of the sample.

Overall, our results appear as a first empirical validation, with reference to European countries, of the idea that fiscal imbalances do impose some additional constraint on the manoeuvrability of monetary policies. This message appears even more relevant in the context of a unified European monetary policy process. Calls for fiscal restraint and sounder national budget policies within the Stability and Growth Pact are addressed on a recurrent basis by the new monetary authorities. Some observers even attribute the current alleged weakness of the external value of the euro to single countries' social security long-term solvency problems.

In the EMS context, the relationships between a country's budgetary position and interest rate determination were somehow made more evident by the further presence of a binding constraint on economic policies –the ERM band. The adoption of exchange rate agreements made financial markets, and the public in general, more aware of the new trade-off faced by monetary authorities, by attaching extra penalties to departures of interest rate policies from an anti-inflationary stance. It is clear that financial markets and policy observers for the very same reason scrupulously and critically monitor the behaviour of the European Central Bank. Whether the interesting results found in this study are going to be observed for the euro area as a whole and for other economies, is left as a direction for further research.

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## Data Appendix

Variables were taken from OECD Main Economic Indicators and IMF International Financial Statistics. In most cases, we were able to employ seasonally adjusted data. For each country we measured real output using the GDP at constant price series. The inflation series were defined as simple 4-quarter log-differences in the all-items CPI. Below we briefly list the short-term interest rates we chose as policy indicators, and the definition of variables in the graphs contained in the Data Appendix. Rates are generally converted from monthly series.

| ***Country*** | ***Modelled Interest Rate Variable*** |
|---|---|
| *FRANCE* | Call Money Rate |
| *ITALY* | 3-Month Interbank Deposits (Overnight) |
| *BELGIUM* | Call Money Rate |
| *IRELAND* | Call Money Rate |

| ***Variable*** | ***Definition*** |
|---|---|
| *EXPINF* | Expected inflation, as described in the main text |
| *ADJSPREAD* | Adjusted spread, as described in the main text |
| *GERFIBOR* | 3-month German Fibor |
| *RESERVES* | 4-quarter log-difference in official reserves excluding gold |
| *M1(3)GROWTH* | 4-quarter log-difference in M1(M3) Growth |

| | *Constant* | *Coefficient* | $R^2$ |
|---|---|---|---|
| ***France*** | *0.025337* *(0.0028092)* | *0.65963* *(0.174470)* | *0.171613* |
| ***Italy*** | *0.030986* *(0.0039753)* | *0.40003* *(0.047464)* | *0.510899* |
| ***Belgium*** | *0.0067862* *(0.0023722)* | *0.39122* *(0.056233)* | *0.41227* |
| ***Ireland*** | *0.032671* *(0.0030398)* | *0.39169* *(0.10448)* | *0.169222* |

**Table 1. Preliminary regressions. France, Italy, Belgium, Ireland, 1980Q1-1997Q2.**
*Results are from RLS regressions of the spread between the yields on national long-term bonds and that on analogous German Bunds on expected inflation (standard errors in parentheses).*

| ***Country/Regressor*** | ***France*** | | ***Italy*** | |
|---|---|---|---|---|
| ***Constant*** | *0.0324*<br>*(0.004901)* | | *0.06893*<br>*(0.007694)* | |
| ***Expected Inflation*** | *0.5583*<br>*(0.115)* | | *0.7067*<br>*(0.07481)* | |
| ***Output Gap*** | *0.152*<br>*(0.1331)* | | *-0.1584*<br>*(0.3602)* | |
| ***GerFibor*** | *0.8559*<br>*(0.06993)* | | *0.2047*<br>*(0.1252)* | |
| ***AdjSpread*** | *0.8625*<br>*(0.07898)* | | *0.9645*<br>*(0.1467)* | |
| *Variable Addition Tests* | *M3 Growth*<br>*ΔReserves* | *0.08616 (0.06205)*<br>*-0.01434 (0.0072)* | *M1 Growth*<br>*ΔReserves* | *0.1347 (0.05864)*<br>*-0.0151 (0.00875)* |
| *Summary Statistics* | $R^2$<br>$\sigma$<br>*DW*<br>*AR 1- 5 F( 5, 60)*<br>*ARCH 4 F( 4, 57)*<br>*Normality* $\chi^2$<br>*RESET* | *0.918652*<br>*0.0101441*<br>*1.81*<br>*0.53061 [0.7522]*<br>*3.5159 [0.0124]*<br>*23.248 [0.0000]*<br>*1.7875 [0.1860]* | $R^2$<br>$\sigma$<br>*DW*<br>*AR 1- 5 F( 5, 60)*<br>*ARCH 4 F( 4, 57)*<br>*Normality* $\chi^2$<br>*RESET* | *0.953379*<br>*0.00831553*<br>*1.45*<br>*1.7068 [0.1469]*<br>*0.81201 [0.5227]*<br>*16.113 [0.0003]*<br>*7.646 [0.0074]* |

**Table 2. Estimated interest rate reaction functions. France and Italy, 1980Q1-1997Q2. Static Long-Run Solutions.**

*All results are obtained from Recursive Least Squares regressions of the monetary instrument on a constant, the indicated regressors, and one lag of the dependent variable. Regressors are defined in the main text. Asymptotic standard errors in parentheses. We tested for the addition of other regressors. Zero restrictions on lagged money growth and the 4-quarter change in the (log of) official reserves of foreign currency were tested by including them in the baseline regression. Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P values in brackets.*

| *Country/Regressor* | *Belgium* | | *Ireland* | |
|---|---|---|---|---|
| ***Constant*** | 0.01362 (0.003863) | | 0.04562 (0.01502) | |
| ***Expected Inflation*** | 0.1441 (0.06617) | | 0.5312 (0.1864) | |
| ***Output Gap*** | -0.02964 (0.05137) | | 0.1962 (0.2266) | |
| ***GerFibor*** | 0.877 (0.06842) | | 0.9121 (0.2122) | |
| ***AdjSpread*** | 0.8809 (0.1076) | | 1.074 (0.2172) | |
| *Variable Addition Tests* | *M3 Growth* | 0.02811 (0.02532) | *M3 Growth* | -0.005084 (0.05689) |
| | *ΔReserves* | -0.01522 (0.00650) | *ΔReserves* | -0.07666 (0.02123) |
| *Summary Statistics* | $R^2$ | 0.942585 | $R^2$ | 0.670091 |
| | $\sigma$ | 0.00641402 | $\sigma$ | 0.0303122 |
| | DW | 2.09 | DW | 1.92 |
| | AR 1- 5 $F(5, 60)$ | 0.54562 [0.7409] | AR 1- 5 $F(5, 60)$ | 0.53245 [0.7508] |
| | ARCH 4 $F(4, 57)$ | 0.30798 [0.8715] | ARCH 4 $F(4, 57)$ | 0.064094 [0.9922] |
| | Normality $\chi^2$ | 21.288 [0.0000] | Normality $\chi^2$ | 128.99 [0.0000] |
| | RESET | 3.7747 [0.0565] | RESET | 2.1777 [0.1449] |

**Table 3. Estimated interest rate reaction functions. Ireland and Belgium, 1980Q1-1997Q2. Static Long-Run Solutions.**

*All results are obtained from Recursive Least Squares regressions of the monetary instrument on a constant, the indicated regressors, and one lag of the dependent variable. Regressors are defined in the main text. Asymptotic standard errors in parentheses. We tested for the addition of other regressors. Zero restrictions on lagged money growth and the 4-quarter change in the (log of) official reserves of foreign currency were tested by including them in the baseline regression. Asymptotic standard errors are in parentheses. AR is a LM test for the hypothesis of no serial correlation; ARCH checks whether residuals have an ARCH structure, with no ARCH as the null; Normality tests the normality of residuals; RESET tests the null of no functional mis-specification. P-values in brackets.*

## Recursive Graphs and Stability Tests

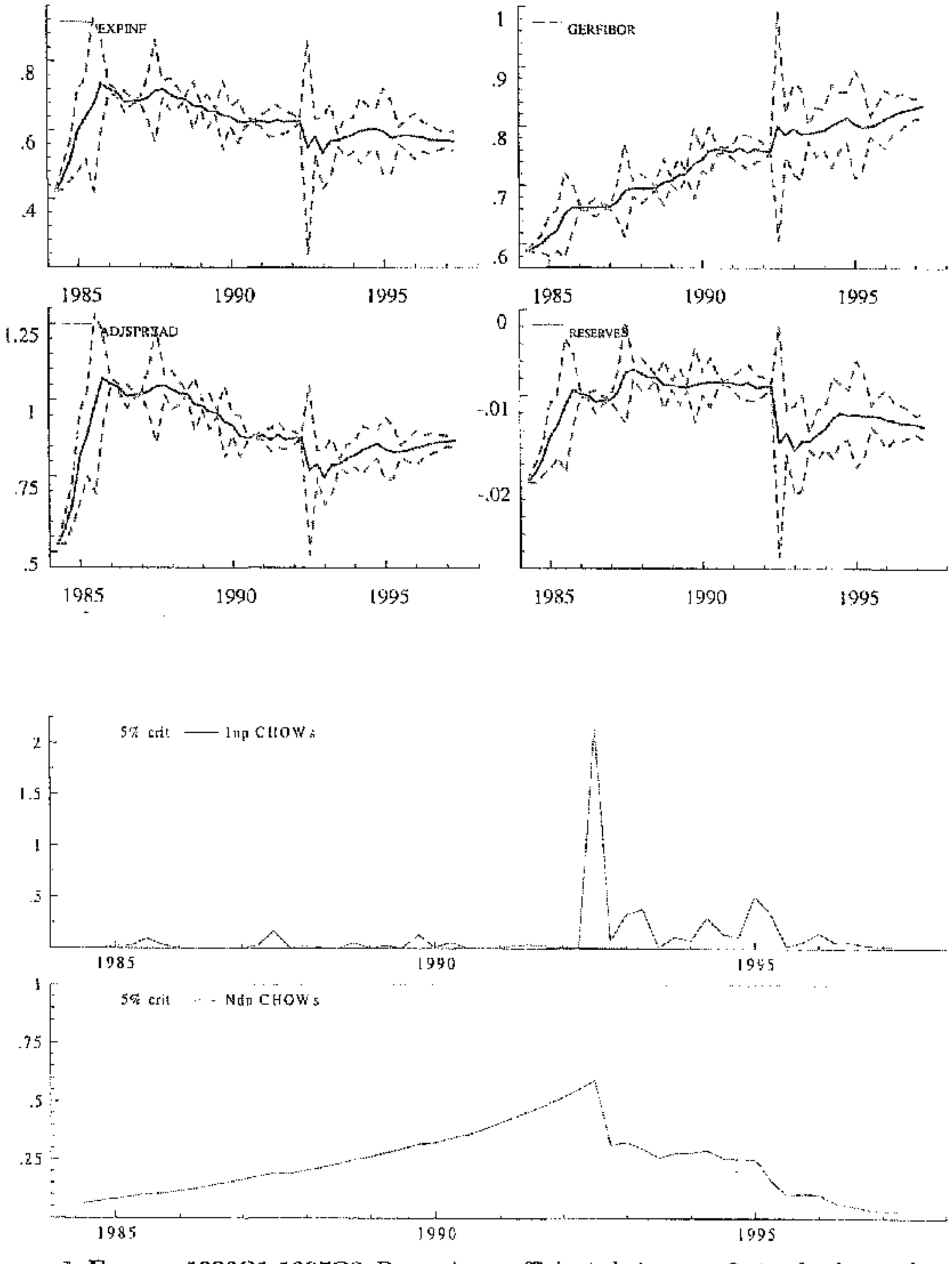


**Figure 1. France, 1980Q1-1997Q2.** Recursive coefficients between ± 2 standard-error bands; 1-step up and N-step down Chow tests (5%).

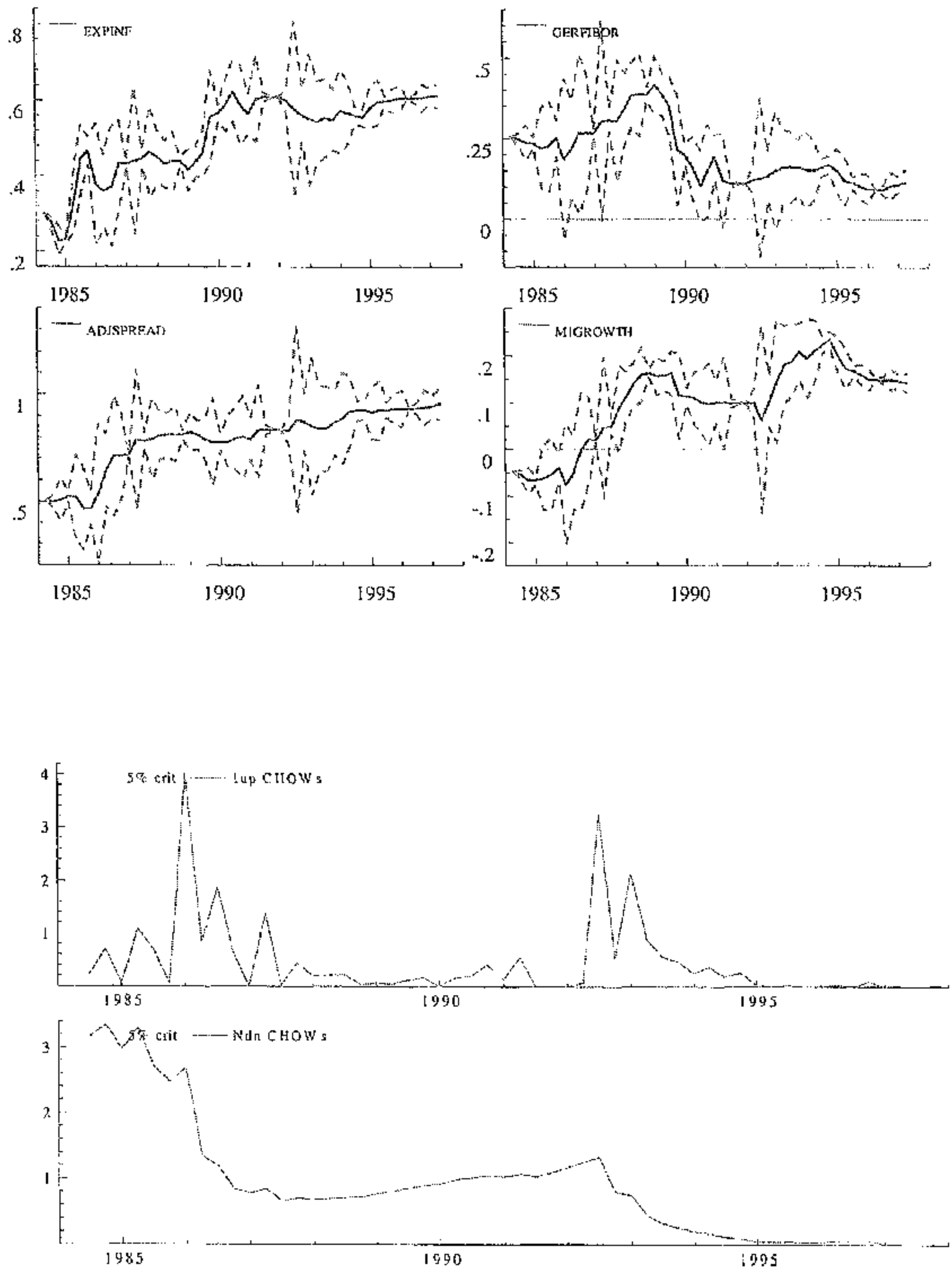


**Figure 2. Italy, 1980Q1-1997Q2.** Recursive coefficients between ± 2 standard-error bands; 1-step up and N-step down Chow tests (5%).

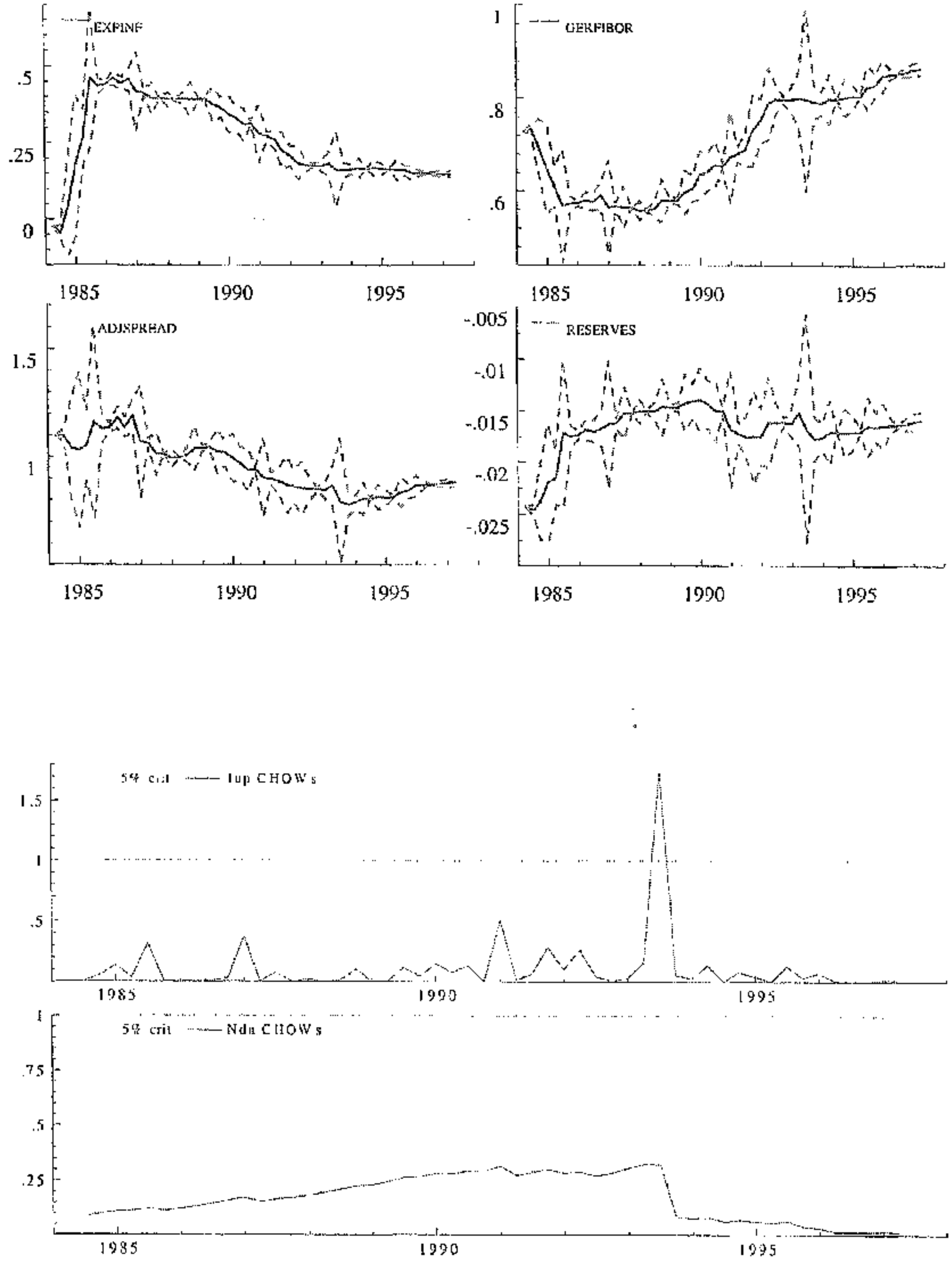


**Figure 3. Belgium, 1980Q1-1997Q2.** Recursive coefficients between ± 2 standard-error bands; 1-step up and N-step down Chow tests (5%).

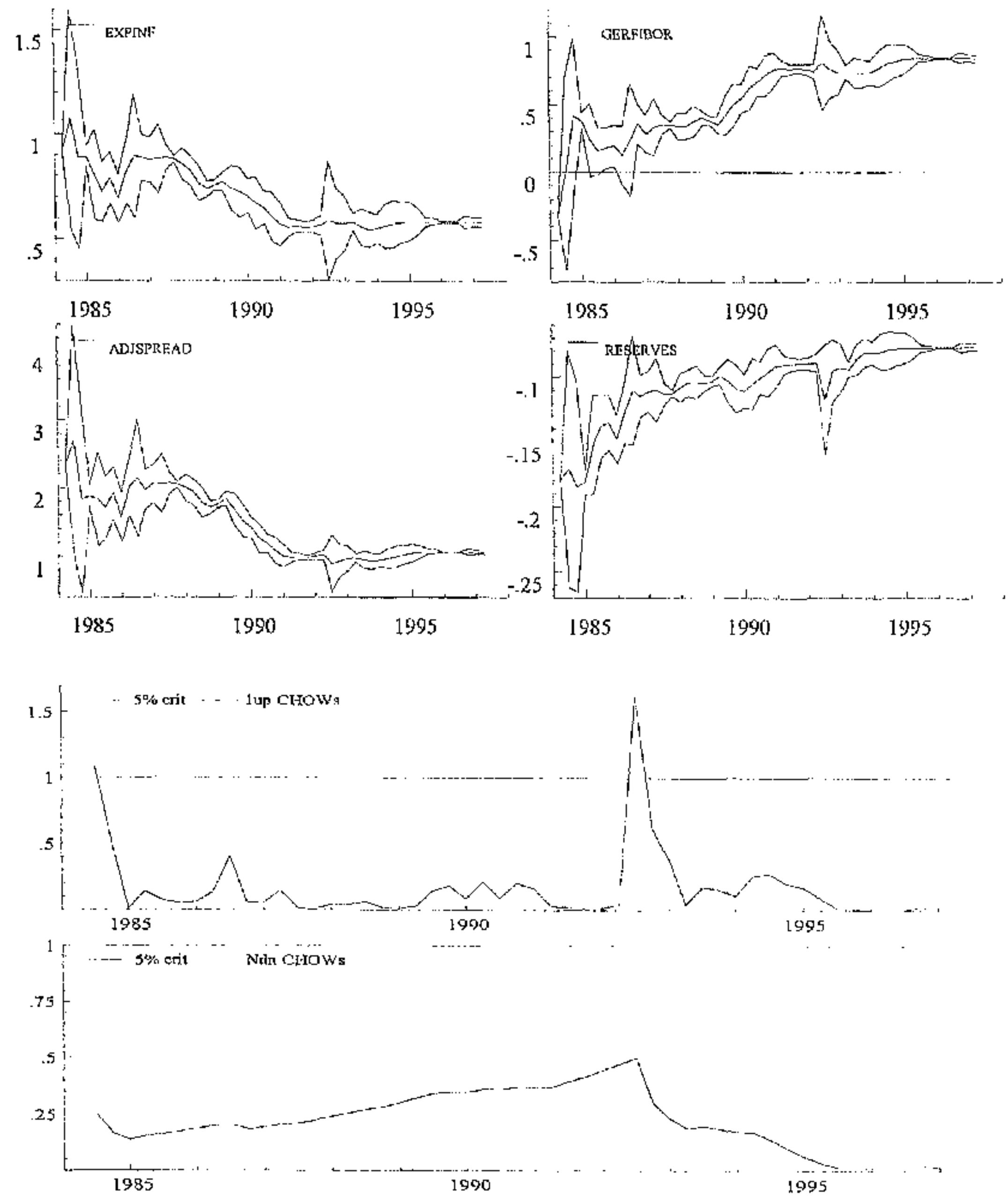


**Figure 4. Ireland, 1980Q1-1997Q2.** Recursive coefficients between ± 2 standard-error bands; 1-step up and N-step down Chow tests (5%).

# The Information Content of Euro Area Monetary Aggregates: Is M3 a Leading Indicator of Inflation Developments?

# Chapter 3

*"...Making money growth an explicit target of monetary policy, or even using money growth in the role of what the literature has called an 'information variable', makes no sense unless observed fluctuations in money anticipate movements of prices, or output, or whatever constitutes the ultimate objective that monetary policymakers seek to achieve. (What would it mean to exploit an information variable that contains no relevant information? What would be the point in pursuing an intermediate target that is not observably intermediate between the central bank's actions and the intended consequences?) In either case, whether movements in money anticipate movements in prices and/or output is crucial..."*

B. Friedman (1996), p. 138

## 1. Introduction

The monetary policy strategy of the Eurosystem, as illustrated in a number of official declarations of European Central Bank's Governing Council[1], consists of three elements. Firstly, the broadly defined primary objective of the Eurosystem, namely price stability over the medium term, is given a quantitative definition ("...increase in the HICP for the euro area of below 2%"). To achieve the stated goal of the strategy, a prominent role is then assigned to the monitoring of monetary aggregates for the area as a whole. This translates into the announcement, each year, of a reference value for the growth of a broad monetary aggregate. For 1999 and 2000, the announced reference value for the growth of M3 was 4.5%. Finally, a broad assessment of the outlook for, and risks to, price stability, completes a pragmatic framework whereby the information obtained from monetary aggregates is jointly gathered with other derived from a number of economic indicators.

It is then clear how, despite the controversies accompanying and following the choice of the monetary policy strategy of the ESCB, the option adopted marks a substantial departure from standard models of monetary targeting regimes. The "prominent" status assigned by ECB's Governing Council to money, though forcefully stated, by no means coincides with the intermediate-objective role monetary aggregates play in textbook treatments of such regimes. Indeed, a bird's eye view of the overall strategy reveals significant departures even from the most closely followed model of inflation control in Europe, i.e., the Bundesbank. Whereas the latter was attaching to the announced value for M3 growth an explicit role of *target*[2], the ECB clearly defines it as *reference value*, qualifying its relevance as the most important, but not unique, *leading indicator* of price developments. However, the recent literature on monetary policy rules (Bernanke *et al.*, 1999; Taylor, 1998, 1999) and some related empirical evidence (Clarida *et al.*, 1998; see our results in Chapter 1), agree that inflation forecast regimes and monetary targeting regimes behave according to similar observed patterns.

Economic theory suggests that money can play two roles in a monetary policy strategy. In common to both inflation forecast- and standard monetary-targeting regimes, the behaviour of monetary aggregates can be usefully monitored by the Central Bank to obtain information about future inflation. Monetary authorities adopting money growth as an information variable assume that past and current *monetary* developments contain useful information about current and future *price* developments.

On the other hand, in standard regimes of monetary targeting, the money stock is seen to provide for a *nominal anchor* to the whole economic system, and to inflation expectations in particular. Consistently with the view that inflation is, ultimately, a monetary phenomenon, the announcement of a target for the growth of some broad monetary aggregate helps the private sector forming expectations about future nominal variables. The path of inflation expectations is thus "co-ordinated" by monetary authorities towards the adopted definition of price stability. Of course, for such strategy to lead to optimal outcomes, some important criteria have to be fulfilled:

a) A stable demand for the monetary aggregate must exist. This involves that a monotonic relationship between money and prices is identified, and that, in turn, a similar one be found between the intermediate and the final goal of the strategy.
b) A monotonic relationship of the available policy instrument -usually a short-run interest rate- with the monetary aggregate must exist. This is equivalent to require that the money stock is controllable at some horizon through changes in the policy instrument.
c) The commitment of monetary authorities to the strategy and the announcement must be credible.

[1] See, for example, ECB (1999a). Angeloni et al. (1999) extensively discuss the motivations of such strategy.
[2] See von Hagen (1999) for an interesting reconstruction of the foundations and development of Bundesbank's strategy since early 1970s.

d) Finally, relationships in both a) and b) must be invariant not only to changes in the behavioural relationships within the economic system, but also to Central Bank's own actions. In other words, the strategy must be immune to Lucas' critique.

It is thus clear that, for a monetary policy strategy aiming at using monetary aggregates, either as an intermediate target (nominal-anchor role) or as an information variable (leading-indicator role), the identification of the statistical properties of the money-prices relationships is crucial.

Well before the start of Stage Three of the European Monetary Union, a number of empirical contributions have then addressed some of the above issues, from a number of perspectives. A first strand of the literature has focused on the estimation of structural models for the euro area, to evaluate their restricted reduced forms in terms of stability and controllability of the M3 stock (Cabrero *et al.*, 1998; Fase and Winder, 1999; Vlaar and Schuberth, 1999; Coenen and Vega, 1999). The results point to positive findings for the former feature, but to a negative response as regards the latter.

More recently, the forecasting properties of euro-area M3 have started to be analysed directly in terms of price developments. This is the case of a recent contribution by Gerlach and Svensson (1999)[3]. This paper assesses the informational content of the output gap (actual output minus its potential level) and the interest rate spread (long-term interest rate minus short-term) along with that of the real money gap (actual real M3 holdings minus their long-run equilibrium value). The exercise, performed within the so-called *Pstar* framework, finds that both the output gap and the real money gap contain useful information about future price developments[4].

A further direction for research is suggested by Sims (1972). This pioneering study points out that investigating the leading indicator properties of monetary aggregates for future price developments is equivalent to studying the Granger-causality properties of money on prices. More precisely, Granger causality from money to prices is the necessary and sufficient condition for money to be a leading indicator of price developments[5].

[3] See also Svensson and Woodford (1999) for a further attempt to fully rationalise the use of monetary indicators for interest-rate policy.

[4] A somewhat deeper discussion of Gerlach and Svensson's (1999) results and *Pstar* models is carried out below.

[5] However, this is neither necessary nor sufficient to guarantee a nominal-anchor role for it. Sims (1972) also shows that the direction of Granger causality has to run from prices to money for the latter to be effectively employed as a nominal anchor. This is in open conflict with the assumption in Estrella and Mishkin (1997), who instead rely on a minimal information content of monetary aggregates as a for money to be used as a nominal anchor.

Consequently, such properties do not necessarily need to be investigated from the restricted reduced form of a structural model: standard Granger non-causality tests are usually performed straight from the unrestricted reduced form of the same model. This result is crucial, for it allows to overcome a number of important hurdles one would meet when trying to identify (or over-identify) a minimally well specified structural model of the euro area.

However, some relevant caveats are in order. First, when information from monetary indicators is to be employed for policy purposes, a clear understanding of the transmission mechanism of policy actions is as relevant as that of the existence of a relationship between money and prices. More importantly, the concept of Granger causality is strictly non-operational, since it refers to the entire universe of relevant information characterising the behavioural relationships within the economic system. Again, this problem is particularly severe in the case of policy-related variables. Inference on Granger causality may not be invariant to the inclusion or exclusion of additional information, i.e., the extension or reduction of the information set. Thus, conclusions about the nature of monetary shocks cannot be firmly drawn based on the limited vector of variables typically included in structural models of the economy.

Bearing these constraints in mind, we devote the first part of this analysis to conduct a series of tests on the null hypothesis that money does not Granger-cause prices in the euro area. The study is carried out within the context of a cointegrated VAR system already used by Coenen and Vega (1999) to estimate a demand function for euro-area M3. Subsequently, the leading indicator properties of M3 are investigated within the *Pstar* framework, in line with the previous attempt by Gerlach and Svensson (1999). In other words, we try to assess the information content, and more specifically the predictive power, of the real money gap, but also of the output gap and a measure of the term spread, for future inflation. Finally, with an eye to the results of previous sections, we perform a number of forecast encompassing tests, to compare the predictive ability of our model *vis-à-vis* that of an empirical model of euro area inflation (Fagan, Henry and Mestre, 2000) in which developments in the money markets do not play any role in forecasting inflation.

The main conclusions of our analysis are as follows. First, there is very little empirical evidence against the null of Granger non-causality of M3 aggregates on prices for the euro area. This finding appears to be robust to a number of crucial changes in the specification of the system estimation, the sample used, and the assumptions about the order of integration of the variables. Second, our forecasting model shows the existence of a reasonably strong positive association between the real money gap and future inflation up to five-six quarters ahead. However, the output gap and, to a lesser extent, the real interest rate or the term spread, tend to display similar predictive ability. This is in line with recent findings for the US[6] (Tzavalis and Wickens, 1996; Stock and Watson, 1999), and allows to conclude that standard *Pstar* models are likely to forgo the rich information content of variables other than the real money gap. Finally, our forecast encompassing exercises suggest that information obtained from "monetary" models like ours and that from more "structural" models of the euro area should be systematically and jointly used to study future price developments.

## 2. Money-Prices Relationships in the Euro Area and Granger-causality tests

In this section we focus on the statistical properties of the relationship between monetary aggregates and inflation as displayed by the results of tests on Granger non-causality conducted within the same framework as in Coenen and Vega (1999). Our aim is to investigate the causality relationships between money and prices within the information set at hand. In the next section, we will show that the causality links found below, while robust to a reduction of this dataset, are not invariant to its extension. In particular, the results of this section will not appear robust to the inclusion, as in Gerlach and Svensson (1999), of variables accounting for monetary policy's implicit inflation objective and for the area-wide output gap.

The series we use (see Data Appendix) come from the European Central Bank Area-Wide Database[7], and are seasonally adjusted quarterly observations from 1980Q1 to 1998Q4, computed using fixed weights based on 1995 GDP at PPP rates. Quarterly averages of M3 monthly data were employed for the broad money aggregate, whereas GDP and the GDP deflator were used for the scale and the price variables. The real money growth measure is the quarterly change in the real stock: $\Delta\tilde{m}_t = \Delta\left(m_t - p_t\right) = \tilde{m}_t - \tilde{m}_{t-1}$,

[6] Friedman (1997) carefully examines a number of theoretical and empirical issues related to the use of monetary indicators for the US interest rate policy.

[7] The series employed in this study are now publicly available. Similar data are also available from the Bank of International Settlements. The February 1999 ECB Monthly Bulletin released historical estimates of these aggregates, while ECB (1999b) contains further information about the identification of the links between the different categories of

whereas the short- and long-term interest rates are computed as averages of national rates weighted using the same weights as above[8].

Let $Z' = (m\text{-}p,\ \pi,\ y,\ R^s,\ R^l) = (m\text{-}p,\ \pi,\ X')$ be a vector of CI(1,1) variables consisting of real holdings of M3 ($m$-$p$), the inflation rate (as measured by the annualised quarterly changes in the GDP deflator, $\pi = 4\Delta p$), real GDP ($y$) and the short- ($R^s$) and long-term ($R^l$) interest rates. The problems inherent to the order of integration of these variables have already been examined, and they are fully discussed in Coenen and Vega[9] (1999). Consider the following VAR model:

$$Z_t = \sum_{i=1}^{k} J^i Z_{t-i} + V_t, \qquad [1]$$

and its VECM representation:

$$\Delta Z_t = \sum_{i=1}^{k-1} A^i \Delta Z_{t-i} + \Gamma \Lambda' Z_{t-1} + V_t, \qquad [2]$$

with Γ and Λ (5, r) full-rank matrices.

Coenen and Vega (1999) estimated model [2] with lag-length $k = 2$ and cointegration rank $r = 3$. In our context, we investigate the leading indicator properties of M3 by performing a battery of tests on the null hypothesis that money *does not* Granger-cause prices. Following the suggestions in Toda and Phillips (1993, 1994), we define the following partitioned matrices:

monetary aggregates and financial instruments included in the old national monetary statistics and the harmonised categories of Monetary and Financial Institutions of the Euro Area.

[8] Obviously, there are non-trivial issues related to the DGP of our aggregate variables. Indeed, the very existence of an area-wide business cycle, monetary policy, financial market, etc. appears to be the key problem for any analysis dealing with similar aggregate issues. However, we feel that some empirical exercise on the objective of our investigation is worth trying well before reliable and uncontroversial statistics for Euroland become available. A more extensive discussion of the weighting procedures and aggregation methods can be found in Coenen and Vega (1999), who also show how their results -and likely ours- are substantially invariant to changes in the aggregation methods.

[9] See also footnote 10 below.

$$
Z_t = \begin{pmatrix} m_t - p_t \\ \pi_t \\ X_t \end{pmatrix} \; ; \; A^i = \begin{pmatrix} A_{11}^i & A_{12}^i & A_{13}^i \\ A_{21}^i & A_{22}^i & A_{23}^i \\ A_{31}^i & A_{32}^i & A_{33}^i \end{pmatrix} \quad i = 1, \ldots, k\text{-}1 \; ;
$$

$$
\Gamma = \begin{pmatrix} \Gamma_1 \\ \Gamma_2 \\ \Gamma_3 \end{pmatrix} \text{ and } \mathrm{A} = \begin{pmatrix} \mathrm{A}_1 \\ \mathrm{A}_2 \\ \mathrm{A}_3 \end{pmatrix} \tag{3}
$$

This way, we are able to impose a series of hypotheses about causality, exclusion and weak exogeneity, both on short- and long-run parameter matrices. These different hypotheses are defined as follows [see Toda and Phillips (1994) for the asymptotic distributions of the corresponding Wald tests]:

$$
\begin{aligned}
&\mathrm{H}^*\!: \; \mathrm{A}_{21}^1 = \ldots = \mathrm{A}_{21}^{k-1} = 0 \text{ and } \Gamma_2 \mathrm{A}_1' = 0; \; \text{(short- and long-run causality)} \\
&\mathrm{H}_\perp^*\!: \; \mathrm{A}_{21}^1 = \ldots = \mathrm{A}_{21}^{k-1} = 0; \; \text{(short-run causality)} \\
&\mathrm{H}_1^*\!: \; \mathrm{A}_1 = 0; \; \text{(long-run exclusion)} \\
&\mathrm{H}_2^*\!: \; \Gamma_2 = 0; \; \text{(weak exogeneity)} \\
&\mathrm{H}_{21}^*\!: \; \Gamma_2 \mathrm{A}_1' = 0; \; \text{(long-run causality)}
\end{aligned} \tag{4}
$$

Toda and Phillips (1993, 1994) devise three sequential causality Wald-type tests to test the causal effects of one variable on another group of variables and viceversa. In particular, the recommended sequence is as follows:

$$
\text{(P1) Test } H_2^* \begin{cases} \text{if } H_2^* \text{ is rejected, test } H^* \\ \text{otherwise, test } H_\perp^* \end{cases}
$$

$$
\text{(P2) Test } H_1^* \begin{cases} \text{if } H_1^* \text{ is rejected, test } H^* \\ \text{otherwise, test } H_\perp^* \end{cases}
$$

$$
\text{(P3) Test } H_\perp^* \begin{cases} \text{if } H_\perp^* \text{ is rejected, reject the null} \\ \text{otherwise, test } H_2^* \text{ and } H_1^* \begin{cases} \text{if both are rejected, test} \\ \; H_{21}^* \text{ if } r > 1, \text{ or reject} \\ \; \text{the null if } r = 1 \\ \text{otherwise, accept the null} \\ \; \text{of noncausality} \end{cases} \end{cases}
$$

After estimation of model [2], we obtain the following results (p-values in brackets):

$$\mathrm{H}^{*}:\ \chi_2^2 = .251\ \ (.882)$$
$$\mathrm{H}_{\perp}^{*}:\ \chi_1^2 = .069\ \ (.793)$$
$$\mathrm{H}_{1}^{*}:\ \chi_3^2 = 12.72\ \ (.005)$$
$$\mathrm{H}_{2}^{*}:\ \chi_3^2 = 26.43\ \ (.000)$$
$$\mathrm{H}_{21}^{*}:\ \chi_1^2 = .251\ \ (.617)$$

Therefore, based on the Toda and Phillips tests conducted above, the null hypothesis that M3 money does not Granger-cause GDP prices within the vector of variables at hand cannot be rejected at standard confidence levels. This finding is also robust to the choice of the sample period. When our baseline VAR is estimated recursively over the sample 1993:Q1 to 1998:Q4, the maximal test statistics for the hypotheses H*, $\mathrm{H}_{\perp}$, and $\mathrm{H}_{21}$* are, respectively: 0.477, 0.268 and 0.937. All of these are well below the corresponding $\chi^2$ critical values[10]. The same results hold if the estimation sample starts in 1985, dropping the first five years from the analysis.

The Phillips-Toda tests conducted above heavily rest on the results of standard pre-tests for unit roots and cointegration[11]. Granger-causality tests conditional on the estimation of VAR parameters may be crucially impaired by potential biases arising from inaccurate determination of the number of unit roots and cointegration rank in small samples like ours. To counter-check our conclusion about the failure to reject Granger non-causality of M3 on prices, we thus study whether it can be attributed to mis-specification of the cointegration rank, and/or of the order of integration of the variables included in our vector.

[10] The use of $\chi^2$ critical values in a recursive context tends to bias results towards rejection of non-causality, since no allowance is made for the endogenous search. In this sense, our conclusion on non-Granger causality appears to be even more firmly grounded.

[11] The application of ADF tests to our vector of variables yields results absolutely in line with the findings of the Johansen cointegration tests performed in Coenen and Vega (1999). Real money balances, income and both interest rates are I(1) for the sample under investigation. As regards $m$ and $p$, however, the evidence is more mixed, and appears not to be invariant to the nature of the trend contained in the alternative hypothesis, and to the sample under consideration. Comforted by results of recursive computation of these statistics, we conclude that the evidence is broadly in favour of the inflation rate and nominal money growth being non-stationary processes for the sample we adopt. Detailed results from these tests are available from the author upon request.

To this end, we follow Toda and Yamamoto (1995), who show how the estimation of VARs in levels and the inference on general restrictions can be efficiently conducted even if the time series involved are integrated or cointegrated of an arbitrary order.

First, we apply the standard lag-selection procedures to determine the lag length ($k^*$) in the VAR. Subsequently, a ($k^*+d_{max}$)th-order VAR is estimated, where $d_{max}$ is the maximum order of integration suspected to occur among the variables in the system (in our case 1 or 2). Finally, we impose and test restrictions on the first $k^*$ coefficient matrices, discarding the last $d_{max}$ lagged terms, as suggested by Toda and Yamamoto (1995).

**Table 1** (left panel) reports results from tests for Granger non-causality of M3 on $p$ in VARs including two sets of variables: $Z^1 = (m, p, y, R^s, R^l)$ and $Z^2 = (m\text{-}p, \Delta p, y, R^s, R^l)$. This way, we are able to develop our tests on alternative assumptions concerning the order of integration of the price series, and the long-run homogeneity of $m$ and $p$. The maximum order of integration is presumed to be 2 (I(2) model) or 1 (I(1) model) in $Z^1$ and 1 (I(1) model) in $Z^2$, allowing for the possibility of nominal variables being I(2) but constraining real variables to be, at most, I(1). In other words, we allow for the possibility of both nominal and real variables being I(1). Accordingly, columns 1 and 2 in the table refer to the specification in $Z^1$, under alternative hypotheses on the maximum degree of integration among the variables included in the vector: $d_{max} = 2$ (col. 1) or $d_{max}=1$ (col. 2). Column 3 refers to the specification in $Z^2$, under the hypothesis $d_{max}=1$. The right panel of table 1 reports tests for model reduction from $Z^1$ to $Z^2$; in other words, we also explicitly test for long-run homogeneity of money and prices. Finally, in line with the evidence from both sequential tests for lag exclusion and information criteria, the VAR's orders are set to $k = 2$ in the shaded cells in the table. However, we also report results for $k^*+1$ and $k^*-1$. The complete set of results in the table permit to comfortably gauge the robustness of the conclusions of the Phillips-Toda tests conducted above against alternative assumptions on maximum order of integration ($d_{max}$) and the VAR lag-length ($k^*$).

**Table 2** repeats the tests for different subsets of the variables in the systems: 1) $Z_a^1=(m, p)$ and $Z_a^2=(m\text{-}p, \Delta p)$; 2) $Z_b^1=(m, p, y)$ and $Z_b^2=(m\text{-}p, \Delta p, y)$; 3) $Z_c^1=(m, p, y, R^s)$ and $Z_c^2=(m\text{-}p, \Delta p, y, R^s)$; and 4) $Z_d^1=(m, p, y, R^l\text{-}R^s)$ and $Z_d^2=(m\text{-}p, \Delta p, y, R^l\text{-}R^s)$. The structure of the table is the same as in Table 1, except for the inclusion of the corresponding Phillips-Toda tests in column 4. This in turn allows us to obtain some clues on the relevance of the composition and size of the information set for the nature of the Granger-causality relationships we are testing for.

By looking at tables 1 and 2, a number of clear-cut conclusions can be drawn. The analysis just performed suggests that the empirical evidence for rejecting Granger non-causality of M3 on prices is very thin. Both the Phillips-Toda tests and the Toda-Yamamoto tests fail to reject Granger non-causality of $m$ on $p$ at standard confidence levels. This finding, which is stable throughout the sample chosen, appears also to be robust to a number of alternative assumptions regarding: i) the maximum degree of integration of the variables in the system; ii) the lag-length selected for the VAR; and, iii) the imposition or not of the long-run homogeneity of money and prices[12]. Finally, the results in table 2 also suggest that the above findings are broadly invariant to a reduction of the information set employed. This conclusion, however, cannot be readily extended to the case in which the vector of variables included in the system is broadened, as we will see in the next section.

## 3. The *Pstar* Concept and Leading Indicator Properties of M3: Forecasting Models of Euro-Area Inflation

### *3.1 An Introduction to the Pstar model*

The policy relevance of the *Pstar* concept dates back to its use by a number of influential studies as the analytical basis for the quantitative derivation of annual monetary targets in various countries[13]. Monetary targets were, of course, introduced in Germany only (since 1975; see Neumann, 1997; von Hagen, 1995, 1999), and the Bundesbank appears to be the only monetary authority actively employing a *Pstar*-type framework. Nonetheless, this indicator of price pressures has re-emerged in the context of discussions on the process of monetary unification in Europe (Svensson, 1999c; Gerlach and Svensson, 1999). In fact, an intuitive illustration of the concept can easily be shown as underlying the quantitative definition of the money-growth reference value announced by

[12] Though the homogeneity hypothesis cannot be rejected at conventional confidence levels

[13] The use of *Pstar* as an indicator of price pressures was initially advocated by studies developed mainly within Central Banks. For instance: US, see Hallman, Porter and Small (1991); Germany, Tödter and Reimers (1994); Japan, Bank of Japan (1990). The evident trending behaviour of the velocity of money in many of these countries subsequently posed severe theoretical challenges to its practical implementation. A study of the *Pstar* concept (applied to Germany and the UK) from a cointegration perspective is undertaken by Funke et al. (1997), while Groeneveld (1998) extends the analysis by applying a Kalman-Filter technique to a number of EMS countries.

the Eurosystem as one of the two "pillars" of its own policy strategy (ECB, 1999a). In the words of the President of the ECB:

> *"...The first reference value for monetary growth decided by the Governing Council... is consistent with the maintenance of price stability according to the ESCB's published definition, while allowing for sustainable output growth. It has been derived by assuming that the trend growth rate of real GDP in the euro area is in the range of 2% to 2 1/2% per annum and the velocity of circulation of M3 declines at a trend rate of between 1/2% and 1% each year...."*
> (Introductory Statements by W. Duisenberg, Press Conference of December 1, 1998)

The basic idea of the *Pstar* approach is to define the equilibrium price level ($p^*$) as current money holdings per unit of potential output in correspondence of an equilibrium level of velocity. From the simple equation of exchange ($p_t \equiv m_t - y_t + v_t$, in logs):

$$p_t^* \equiv m_t - y_t^* + v_t^* \qquad [5]$$

Thus, deviations of the current price level from this long-run equilibrium can only be possible if current output and/or velocity depart from their respective equilibrium values

$$p_t - p_t^* = \left(y_t^* - y_t\right) + \left(v_v - v_t^*\right), \qquad [6]$$

and may affect inflation according to some law of motion. If economic activity exceeds its potential value, and/or a expansion in liquidity causes velocity to exceed its long-run level, inflationary pressures will therefore develop. The adjustment between inflation and the price gap may follow, for instance,

$$\pi_t = \Delta p^* - \gamma\left(p_{t-1} - p_{t-1}^*\right) + z_t, \qquad [7]$$

where $z_t$ is a white-noise disturbance. Thus, if the price gap in the previous period was indeed positive ( $p_{t-1} - p^*_{t-1} > 0$ ), current inflation would be rising to eliminate the price gap. By contrast, and assuming there is a long-run stable relationship between the two variables[14], if $p_{t-1} = p^*_{t-1}$ both price levels will grow at the same rate. In this instance, if the monetary authorities aim at stabilising inflation around a pre-set target $\tilde{\pi}_t$, the money growth value consistent with such objective would simply be:

$$\Delta m_t = \Delta y^*_t - \Delta v^*_t + \tilde{\pi}_t , \qquad [8]$$

which fully explains the rationale behind ESCB's choice for M3 reference value. It is interesting to note that according to this framework, an interest rate policy *exclusively* geared at targeting money growth is in fact targeting inflation, and it does so *explicitly*[15].

Despite its widespread and influential use as an empirical tool for monetary policy, the *Pstar* framework has no solid theoretical background. Any work that aims at providing some evidence on its empirical properties faces the hard task of setting up reduced-form relationships with no microfoundations at hand. More importantly, this set-up suffers from some of the well-known drawbacks associated with the applications of the quantity theory of money to policy variables over a short- and medium term-horizon. That is, the velocity of circulation of money is in practice endogenous with respect to output, inflation, and the situation of money markets (liquidity, interest rates, etc.). It is then inappropriate to infer causality and forecasting properties from a model that does not take into consideration that velocity is simultaneously determined alongside other real and nominal variables. Finally, the very existence of a velocity trend is difficult to rationalise in the context of a model that has virtually no microfoundations. It is then clear how the statistical properties of the M3-prices relationships should be more properly studied by nesting the *Pstar* concept within alternative models of price pressures, relying on more "structural" frameworks than the one just sketched.

[14] Groeneveld (1998) and Funke et al. (1997) provide some evidence concerning such point

### 3.2 *Structural Models of Money in Europe*

Amongst the first to investigate the money-prices relationships in the euro area, Vlaar and Schuberth (1999) estimate a structural VECM including inflation, real income, short- and long-term interest rates and financial wealth, and identify the structural shocks within the model. Closer to our purposes, the study derives impulse responses that show how (proxied) deviations of real money balances from equilibrium do tend to create some inflationary pressures, as one would expect. Finally, although stable money demand and excess money-inflation relationships are identified, controllability of the money stock via short-term interest rates is rejected.

Coenen and Vega (1999) estimate a demand function for M3 in the euro area spanning 1980-1998. System estimates conducted with the Johansen (1995) procedure and applied to a vector comprising real money holdings, income, long and short interest rates and inflation led to the identification of three cointegrating vectors. The spread between the long and short interest rate, a Fisher-type long-term interest rate parity, and a relation expressing real M3 holdings as a function of income and the nominal and real stochastic trends driving the overall system are identified. Weak exogeneity of real income, interest rates and inflation with respect to the long-run parameters of the latter cointegrating vector is subsequently not rejected. This involves accepting the idea that M3 does not Granger-cause inflation. In the light of this, a long-run conditional money demand equation is estimated. Furthermore, a dynamic, single-equation model of M3 is developed and tested for weak exogeneity of real income, interest rates and inflation with respect to the short-run parameters, and for stability, confirming the previous results. These findings prove to be robust to alternative aggregation methods. Further investigation into the identification of the structural innovations to the multivariate money demand through impulse response analysis broadly validates Vlaar and Schuberth's (1999) findings about the lack of controllability of the money stock via standard policy instruments.

Closer to the problem at hand, Gerlach and Svensson (1999) nested a *Pstar*-type model of inflation and a simple expectations-augmented Phillips curve:

[15] The majority of estimated interest rate reaction functions for the Bundesbank detect a more significant role for expected inflation than for M3, which empirically corroborates the above statement. See Chapter 1, and Muscatelli et al., 1999; Clarida et al., 1998. Svensson (1999a, b) elaborates more extensively on this point.

$$\pi_t = \pi^e_{t-1} + \delta\left(\tilde{m}_{t-1} - \tilde{m}^*_{t-1}\right) + \delta_2\left(y_{t-1} - y^*_{t-1}\right) + \varsigma_t, \qquad [9]$$

where $\pi^e$ represents inflation expectations and $\tilde{m}$ indicates long-run real money balances. Where starred, the latter are calculated in correspondence of trend output, velocity, and real interest rates.

The rest of Gerlach and Svensson's model is as follows:

$$\pi^e_{t-1} = (1-\alpha_\pi)\hat{\pi}_t + \alpha_\pi \pi_{t-1} \quad [9.1]$$
$$\hat{\pi}_t = \exp(\tau_0 + \tau_1 t) \quad [9.2]$$
$$\tilde{m}_t \equiv (m-p)_t = k_0 + k_y y_t - k_i(R^l_t - R^s_t) \quad [9.3]$$
$$\Delta\tilde{m}_t = \delta_0 - \delta_1[\tilde{m} - k_0 - k_y y + k_i(R^l - R^s)]_{t-1} - \delta_2(\pi_t - \hat{\pi}_t) + \delta_3 \Delta\tilde{m}_{t-1} + \upsilon_t \quad [9.4]$$
$$p^*_t = m_t - k_0 - k_y y^*_t + k_i(R^l - R^s)^* \quad [9.5]$$
$$\tilde{m}_t - \tilde{m}^*_t \equiv (m-p)_t - (m-p^*)_t \equiv -(p_t - p^*_t) = (m-p)_t - k^* - k_y y^*_t \quad k^* = [k_0 - k_i(R^l - R^s)^*] \quad [9.6]$$

Inflation expectations are determined in [9.1] on the basis of past inflation and authorities' implicit inflation objective ($\hat{\pi}_t$). The latter in turn is computed ([9.2]) as a deterministic exponential trend, under the constraint that inflation equalled 1.5% in 1998. This particular assumption, kept also in our study, attempts to provide an elementary rationale to the inflation downward trend observed in Europe since mid-eighties. Though not particular sophisticated, the hypothesis of an exponential trend aims at simplifying the task of devising more detailed explanations of the same phenomenon, while stressing the importance of monetary authorities' increasing commitment to lower inflation in the area.

The definition of real money balances in [9.3] follows the standard treatments of long-run real money balances, assumed to depend positively on real income and negatively on the opportunity cost of holding money (as measured by the term spread)[16]. Equation [9.4] postulates Gerlach and Svensson's modelling of the short-run demand for real money holdings, while [9.5] is similar to our equation [6], with the term spread replacing the deviations of velocity from trend. Finally, [9.6] defines the real money gap as the negative of the price gap.

Assuming that long-run equilibrium is characterised by having

[16] See Ericsson (1999) for an up-to-date analysis of recent findings on the empirical specification of money demand functions

$$y_t = y_t^*; \quad p_t = p_t^* \ (\tilde{m}_t = \tilde{m}_t^*); \ (R_t^l - R_t^s) = (R^l - R^s)^*; \ \pi_t = \pi_{t+1} = \hat{\pi}_{t+1}, \qquad [10]$$

(weighted non-linear LS) system estimates of this model yield (standard errors in parentheses):

$$\pi_t - \pi_{t-1} \equiv \Delta\pi_t = -\ .415\ (\pi_{t-1} - \hat{\pi}_t) + \underset{(0.074)}{.163}\ (\tilde{m} - \tilde{m}^*)_{t-1} + \underset{(0.094)}{.021}\ (y - y^*)_{t-1} + \hat{u}_t \qquad [11]$$

Inflation dynamics is then seen as dependent on the deviations of real money holdings from their equilibrium level and departures of past inflation from authorities' implicit objective[17]. These empirical findings broadly validate the *Pstar* concept, since they confirm that price pressures arise from disequilibria in inflation expectations and/or past liquidity holdings. Although the relevance of the output gap is rejected at conventional confidence levels, this measure of the cycle appears to be capturing price pressures not otherwise explained by disequilibria in the money markets and expectations. We turn to this important point later, because a similar results, but with a much stronger statistical significance, will crop up in our results as well.

One conclusion from Gerlach and Svensson's findings pertains to the results of our Granger-causality analysis. The very presence of the real money gap in [11] implies that money Granger-causes prices once one accounts for the presence of authorities' inflation objective and the output gap in the information set. This in turn validates our early argument that inference on the forecasting properties of M3 is not invariant to the addition of relevant information in the vector of variables under consideration. We now build upon this point by extending the analysis towards more systematic inflation forecasting for the euro area.

[17] Gerlach and Svensson also argue that the Eurosystem's money growth reference value performs worse than alternative indicators of monetary disequilibrium, namely the real money gap as computed in [9]. Noting that $(m - \tilde{m})_t \equiv (m - \tilde{m})_{t-1} + \Delta(m - \tilde{m})_t$, the result in [11] can be easily re-parameterised in terms of the two RHS variables in the above identity, the latter of which could be interpreted as the deviations of money growth from the reference value. It follows that, provided due account is taken of the prevailing liquidity situation as measured by the level gap, $(m - \tilde{m})_{t-1}$, $\Delta(m - \tilde{m})_t$ provides the same information content as $(m - \tilde{m})_t$, partly contradicting Gerlach and Svensson's claim.

## 4. Forecasting Inflation in the Euro Area

### *4.1 The Information Content of Money, Output and the Interest Rates in an Extended Framework*

The task we set ourselves is to obtain a measure of disequilibrium in the money markets that does not depend to any extent on weak exogeneity assumptions, as these proved to be dependent on the information set employed.

Simple extension of Coenen and Vega's (1999) two-step Johansen estimation, to a vector of variables now including the output gap and the authorities' inflation objective, would be cumbersome and inappropriate, for a number of reasons:

- it would make impossible the structural identification of the exact cointegrating relationships;
- simultaneity biases are likely;
- we wish to study M3-prices relationships regardless of weak exogeneity assumptions, in turn necessary to identify long-run money demand parameters.

We then start our analysis by considering the application of Johansen's (1995) procedure to the largest vector of variables we adopted for our Granger-causality tests, namely $Z' = (m\text{-}p,\ \pi,\ y,\ R^s,\ R^l)$. That is, we test for cointegration in a system including real money balances, real income, short- and long-term interest rates, and inflation. This way, we aim at obtaining measures for the long-run income elasticity and the semi-elasticity of inflation with respect to long-run money demand, to be subsequently inserted in our measure of the real money gap. We avoid trying to infer the interest rate semi-elasticities of money demand, as that would involve imposing further, costly restrictions on an already complex system.

For the sample 1980Q1-1998Q4 we are thus able to identify the same number of stable cointegrating vectors as in Coenen and Vega (1999), and more precisely[18]:

$$
\begin{aligned}
ecm1_t &= (m-p)_t - \underset{(0.037)}{1.158}y_t + \underset{(0.189)}{1.278}\pi_t + k \\
ecm2_t &= (R^l - R^s)_t - (R^l - R^s)^* \\
ecm3_t &= (R^s - \pi)_t - (R^s - \pi)^*
\end{aligned}
\qquad [12]
$$

[18] Results on cointegration tests are not shown here for brevity, but are available from the author upon request.

In [12], the second and third vectors represent the spread between long- and short-term interest rates and the real short-term interest rate (consistently with the Fisher parity), respectively, both stationary around constants[19]. The first cointegrating vector, in turn, allows us to unambiguously identify the two money-demand elasticities we were looking for.

All estimates are pretty close to those obtained through Coenen and Vega's solved conditional ADL model for long-run money demand. Subsequently, we estimate a VAR system in $\Delta Z_t$, conditional on $\Delta Z_{t-1}$, $ecm1_{t-1}$, $ecm2_{t-1}$, and $ecm3_{t-1}$. The results we obtain for the inflation equation, after less significant variables are excluded, are as follows (standard errors in parentheses)[20]:

$$\Delta\pi_t = \underset{(.13)}{.45}\ [(R^l - R^s)-(R^l - R^s)^*]_{t-1} + \underset{(.10)}{.54}\ [(R^s - \pi)-(R^s - \pi)^*]_{t-1} + \hat{v}_t \qquad [13]$$

| | | | |
|---|---|---|---|
| *T=73 (1980:Q4-1998:Q4)* | *R2=0.30* | *σ=.88%* | *DW=1.95* |
| *LM(1)=1.34 (.252)* | *LM(4)=1.46 (.231)* | *LM(1,4)=0.60 (.667)* | |
| *ARCH(4)=0.50 (.734)* | *HET=0.80 (.553)* | *NORM=0.01 (.997)* | |
| *RESET=1.33 (.252)* | *FOR(24)=7.36 (.999)* | *CHOW(24)=.31 (.999)* | |

Standard tests for the exclusion of additional dynamics on the two crucial terms $\Delta Z_{t-1}$ and $ecm1_{t-1}$ fail to reject the null ([F(8,63)=0.88 (.540)]), thus we conclude that $\hat{v}_t$ is a pure innovation relative to the information set we employ in [13]. In particular, the test for excluding $ecm1_{t-1}$ yields F(1,70)= 0.25 (.617).

The specification above is similar to what previously estimated by Tzavalis and Wickens (1996) for the US. It is clearly inappropriate to give [13] a policy interpretation. What the estimated equation signals is that the past spread and real short interest rate helps predicting current inflation -though to a limited extent. In particular, one might want to rationalise the positive signs in the above expression using the idea that the term structure

[19] Respectively, about 0.613% and 5.2%

[20] The summary statistics are to be interpreted as follows. *LM(i)* and *LM(1,i)* are the Lagrange multiplier F-tests for residual autocorrelation of order $i^{th}$ and up to the $i^{th}$ order, respectively; *ARCH* is the Engle F-test for autoregressive conditional heteroskedasticity; *HET* is the White F-test for heteroskedasticity; *NORM* is the Doornik and Hansen $\chi^2$-test for normality; *RESET* is the regression specification F-test due to Ramsey; *FOR* and *CHOW* are the out-of-sample forecast test and the Chow test for parameter stability over the period 1993:Q1-1998:Q4.

tends to react quite quickly to changes in expected inflation. One can also add that in correspondence of such reaction the likelihood of anti-inflationary policy actions tends to push the real short-term interest rate above its equilibrium value, even before any action is taken in practice. Finally, since the two estimated coefficients do not significantly differ from one another, one can easily show that $\left(R^s-\pi\right)_t=\left(R^l-\pi\right)_t-\left(R^l-R^s\right)_t$. This in turn would call for an elementary re-parameterisation of [13] in terms of the real (*ex-post*) long-term interest rate and the spread only.

As in Tzavalis and Wickens' main findings, our estimates allow us to conclude that both terms are helpful in predicting future GDP inflation in the area. The explanatory power of the above relationship, though, looks limited, as suggested by the low portion of inflation variability explained by the model. Nonetheless, the estimated equation appears well specified, with tests displaying no signs of severe serial correlation, heteroskedasticity or non-normality. No major problem crops up when the equation is used to produce one-step ahead forecasts over the last six years of the sample. The process of monetary convergence among EMS Member Countries that was then taking place, however, calls for some caution in drawing conclusions about this part of the sample[21]. That is, we need to take into account the marked reduction of exchange rate risk premia over the last part of the sample, when uncertainties about the start of Stage Three of EMU were finally removed. We do so by modelling a step reduction in the equilibrium short-term real interest rate from 5.4% (approximately the weighted average of national ex-post real interest rates in the period 1980:Q1 to 1997:Q7) to 3.0% as from 1998Q1, four quarters ahead of the introduction of the single currency. In the same sense, recursive estimation of [13] produces relatively constant parameters over the recent period. **Figure 1a** records time series of fitted and actual values of $\Delta\pi$, the scaled residuals and the residual correlogram, while **Figure 1b** shows a graphical summary of results from the recursive estimation of equation [13] over the 90's.

With the statistical limitations of our model in mind, we turn to introduce in our baseline specification the money gap, $(\tilde{m}-\tilde{m}^*)$, instead of the cointegrating vector $ecm1_t$.

[21] Gerlach and Schnabel (1999), for example, compute a measure of *credibility adjusted* equilibrium real interest rate for the euro area by regressing the weighted ex-post real rate on the average rate of depreciation of the nominal exchange rate against the Deutsche mark. The intercept of that regression could be interpretable as the equilibrium real rate that would prevail assuming no depreciation vis-à-vis the DM. The resulting estimate is 3.55%, with a standard error of 0.96%.

This re-parameterisation is made possible by the following simple relationship linking the real money gap and the error correction term arising from the estimation of the cointegrating vector for long-run money demand $[ecm1_t \equiv \tilde{m}_t - k_0 - k_y y_t + k_i(R_t^l - R_t^s)]$:

$$ecm1_t \equiv (\tilde{m}-\tilde{m}^*)_t - k_y(y-y^*)_t + k_i[(R^l - R^s)_t - (R^l - R^s)^*] \qquad [14]$$

Next, we adopt a new formulation that encompasses equation [13] and Gerlach and Svensson's specification as described in equation [11]. In other words, we augment our parsimonious initial VAR with the additional terms $(\pi_{t-1} - \hat{\pi}_t)$ and $(y-y^*)_{t-1}$.

The new model for inflation is then the following:

$$\begin{aligned}\Delta\pi_t = &-\theta_0(\pi_{t-1} - \hat{\pi}_t) + \theta_1(\tilde{m}-\tilde{m}^*)_{t-1} + \theta_2(y-y^*)_{t-1} + \\ &+\theta_3[(R^l - R^s) - (R^l - R^s)^*]_{t-1} + \theta_4[(R^s_{t-1} - \hat{\pi}_t) - (R^s - \pi)^*] + \hat{v}_t\end{aligned} \qquad [15]$$

where now the real money gap is defined in correspondence of potential output and the inflation objective

$$(\tilde{m}-\tilde{m}^*)_{t-1} \equiv (m-p)_{t-1} - k^* - 1.158 y^*_{t-1} + 1.278\hat{\pi}_t \qquad [16]$$

The two unobservables in the above equation are derived, for comparison purposes, using the deterministic exponential trend postulated in Gerlach and Svensson (1999) for $\hat{\pi}$ and computing potential output by means of a Hodrick-Prescott filter with smoothing parameter $\lambda = 1600$[22].

[22] The use of two-sided filters to proxy potential GDP is standard in models comprising the output gap [see, for instance, Roberts (1997) in the context of the estimation of a Phillips curve model] due to the lack of clear-cut alternatives. This procedure, along with the one adopted to account for inflation expectations and the time-varying inflation objective, is certainly not exempt from both conceptual and econometric problems. More satisfactory measures would perhaps involve applying some errors-in approach to the orthogonalisation of private sectors' misperceptions about future

Estimation of the extended model over the sample 1980:Q4-1998:Q4 yields the following results for equation [15]:

$$\Delta\pi_t = -\underset{(.106)}{.783}\ (\pi_{t-1}-\hat{\pi}_t) + \underset{(.072)}{.196}\ (\tilde{m}-\tilde{m}^*)_{t-1} + \underset{(.130)}{.262}\ (y-y^*)_{t-1} +$$

$$+ \underset{(.073)}{.249}\ [(R^s_{t-1}-\hat{\pi}_t)-(R^s-\pi)^*] + \hat{v}_t \qquad [17]$$

| | | | |
|---|---|---|---|
| *T=73 (1980:Q4-1998:Q4)* | $R^2$*=0.45* | *σ=.79%* | *DW=1.92* |
| *LM(1)=.72 (.399)* | *LM(4)=.85 (.359)* | *LM(1,4)=0.68 (.607)* | |
| *ARCH(4)=0.14 (.968)* | *HET=.221 (.986)* | *NORM=1.23 (.541)* | |
| *RESET=0.15 (.904)* | *FOR(24)=8.5 (.999)* | *CHOW(24)=.31 (.998)* | |

Once again, when we test for the exclusion of additional dynamics from ($\Delta Z_{t-1}$ and $ecm2_{t-1}$) in [17], we fail to reject the null: [F(6,63) = 1.31 (.315)]. A cursory look at **Figures 2a-2b** confirms the absence of major mis-specifications, and the substantial constancy of recursively-estimated parameters.

Our findings tend to provide some support to the idea that the real money gap has substantive predictive power for future inflation in the euro area. However, contrary to the pure *Pstar* specification, model [17] shows that the real money gap (or the negative of the price gap) is not able, by itself, to fully anticipate future inflation. The presence of additional terms in [17] indicates that real interest rates and the output gap have some predictive power for future inflation, and that information obtained from forecasting the level of economic activity and the term structure may turn to be useful in monitoring price developments. Another, more subtle, feature of the above results comes from the observation that they are obtained through a (modified) error-correction specification for the monetary disequilibrium. Hence, it is explicitly recognised that short-run deviations of real money balances from equilibrium are likely, if not normal. Therefore, using M3 growth as a strictly exclusive indicator for future price developments may prove to be inappropriate[23]. As Stock and Watson's (1999) analysis for the US (whom we refer again

inflation, and some gradual learning process for the generation of the output gap (also applicable in the case of the unobservable stochastic component of inflation). The first chapter of this thesis and Muscatelli, Tirelli, and Trecroci (1999) offer some arguments in favour of using a Kalman-filter procedure to overcome some of these drawbacks. In the context of the present investigation, we however prefer to stick with more mainstream solutions, given the need for encompassing exercises with rival models of inflation forecasting.

[23] Friedman (1997) strongly supports the same view for the US.

shortly) seems to point out, "more structural" indicators of the level of economic activity are likely to be more effective in capturing such developments.

### *4.2 What Leading Indicator(s) for Euro-Area Inflation?*

To draw some tentative conclusion on the relative importance of the various indicators we introduced in predicting future inflation, we now evaluate the correlation at different time horizons between $(\pi-\hat{\pi})_t$ and the various indicators included in [17]. In doing so, we follow the approach taken by Stock and Watson (1999) for the US, by producing a series of forecasting equations for horizons ranging from $h = 1$ to $h = 12$ quarters. The models we estimate are generally articulated as follows:

$$\begin{aligned}(\pi-\hat{\pi})_{t+h-1} = {} & \alpha_1^h(\pi_{t-1}-\hat{\pi}_t)+\alpha_2^h(\tilde{m}-\tilde{m}^*)_{t-1}+ \\ & +\alpha_3^h(y-y^*)_{t-1}+\alpha_4^h[(R^s_{t-1}-\hat{\pi}_t)-(R^s-\pi)^*]+\hat{v}_{t+h-1}\end{aligned} \qquad [18]$$

where $v_{t+h-1}$ follows a MA(h-1) process[24].

**Table 3** displays our results. The most evident feature is that there exists a substantive positive association between the real money gap and future inflation. This significant correlation spans over five-six quarters and reaches a peak at the three-to-four quarter horizon. Again, the real money gap does not appear to be a sufficient statistic for future inflation. The output gap and the difference between the real short-term rate and the estimated equilibrium real short-term rate contain valuable information over and above that already contained in the real money gap.

Moreover, according to our findings, the lead over which there is a positive association between output gap and future inflation may indeed be longer than in the case of the real money gap. It stretches over six quarters and reaches a maximum at the five-quarter horizon. The leading information contained in the real rates is instead relatively short-lived, extending over just three quarters. Finally, the sharp reduction in the $R^2$ of the regressions also shows how the forecasting performance of the model dramatically deteriorates as the time horizon increases. Beyond the sixth quarter, no indicator analysed

[24] It should be noted that for $h=1$, [18] is just a trivial re-parameterisation of equation [17]. Equation [18] is often interpreted as a forecasting equation. See, for instance, Clements and Hendry (1996a, b).

appears to contain helpful information. **Figure 3** contributes to further stress this result, by showing inflation deviations from the computed target along with the various (lagged) indicators entering our forecasting equations.

We finally perform some robustness checks on our findings. In particular, we ask ourselves whether the regularities we detect are robust with respect to the specification of added regressors, namely the output gap and the implicit inflation objective. In other terms, we investigate whether altering the way in which our *ad hoc* measures of the unobservables present in the forecasting equations has some bearing on the results we obtain.

**Table 4** reproduces the estimates of model [15], this time allowing, in turn, for alternative methods of generating the potential output and the implicit inflation objective series. In particular, the left panel of the Table maintains the latter as computed through the usual deterministic exponential trend, while allowing potential output to be measured using a) the specification in the European Central Bank Area-Wide Model[25], and b) different smoothing parameters for the HP filter. The right-hand panel, conversely, keeps the baseline HP ($\lambda = 1600$) specification for potential output, while allowing for various HP-based measures of $\hat{\pi}$.

The conclusions to be drawn from results in Table 4 are the following. First, the substantially good leading indicator properties of the real money gap we devised appear to be robust to alternative empirical measures of potential output employed. The same applies, though to a lesser extent, when different empirical measures of monetary authorities' inflation objective are employed. Second, the relevance of the output gap is to vary crucially with its own empirical measurement. The smoother the filter employed for deriving potential output, the less significant the output gap becomes in our baseline model. Finally, the real interest rates and -to a lesser extent- the term spread are confirmed as containing valuable information for future inflation. Such information appears to extend over and above that already contained in the real money gap and the output gap. This finding also appears quite robust across alternative specifications for the unobservable measures of potential output and the inflation objective.

[25] Potential output in the AWM is estimated from a Cobb-Douglas production function with smoothed Solow residuals

## 5. Monetary or "Structural" Indicators? Encompassing Rival Models of Euro-Area Inflation

The main result of the previous sections is the ability of our estimated model for GDP inflation to account for the most salient features of the money-prices relationships in the euro area. Moreover, we evaluated the predictive content of various indicators, and tested for robustness of our results against a number of statistical (homoskedastic innovation errors; constant parameters; information set reduction and extension; etc.) and non-statistical (the alternative specifications for real money gap, output gap and the interest rates magnitudes) criteria. Though the model we specified appears to satisfactorily perform across such checks, we now want to ask ourselves how its forecasts fare against those produced by competing models of inflation in the euro area. Closer to our purposes, in particular we wish to compare the forecasting properties of our model to those of *significantly different* explanations of euro area inflation.

Following Mizon and Richard (1986), and Hendry and Richard (1989), we define a *congruent encompassing* model as a model that is congruent and that is able to account for, or explain, the results obtained by rival models. In this sense, encompassing is a stricter requirement than 'better fit' or (when applied in the context of forecasting) 'lower root mean square forecast error', for it involves that the rival model does not contain any additional information relative to the model at hand.

Fagan, Henry and Mestre (2000) build a somewhat more "structural" model than ours for euro-area GDP inflation. In their work, GDP prices are pinned down in the long-run by trend unit labour costs, which depends in turn on potential GDP and the NAIRU. In the short-run, instead, GDP inflation is a function of changes in trend unit labour costs, changes in import prices and deviations of real trend unit labour cost from equilibrium. It is relative to forecasts produced within their model that we now compare the predictive ability of our own model.

Estimation of the Fagan, Henry and Mestre (2000; FHM, henceforth) model over the sample 1980:Q4-1997:Q4 yields the following results:

$$
\begin{aligned}
\Delta\pi_t = \underset{(.005)}{.021} - \underset{(.109)}{.776}\ \pi_{t-1} + \underset{(.043)}{.140}\ \Delta w^*_t + \underset{(.047)}{.068}\,\Delta w^*_{t-1} + \underset{(.045)}{.136}\ \Delta w^*_{t-2} + \\
+ \underset{(.014)}{.024}\,\Delta pm_{t-1} - \underset{(.076)}{.274}\ (p - w^* - \theta)_{t-1}
\end{aligned}
\qquad [19]
$$

| | | | |
|---|---|---|---|
| $T=69$ (1980:Q4-1997:Q4) | $R^2=0.52$ | $\sigma=.78\%$ | $DW=2.02$ |
| $LM(1)=0.22$ (.641) | $LM(4)=0.23$ (.634) | $LM(1,4)=0.54$ (.709) | |
| $ARCH(4)=0.48$ (.750) | $HET=0.38$ (.964) | $NORM=4.12$ (.127) | |
| $RESET=1.55$ (.218) | $FOR(24)=16.019$ (.887) | $CHOW(24,37)=0.535$ (.945) | |

where $w^*$ and $pm$ stand, respectively, for trend unit labour costs and import prices.

To discriminate between our model (T, henceforth, as described in equation [17]) and FHM, we first perform Mizon and Richard's (1986) *Simplification Encompassing Test* (SET). This way, T and FHM are tested against the so-called minimal nesting model.

In addition, we check the performance of both models against two out-of-sample forecasting tests: the *Forecast-Differential Encompassing Test* (Chong and Hendry, 1981), and the *Forecast-Model Encompassing Test* (Ericsson, 1992). The former involves, *inter alia*, the estimation of parameters $\alpha$ and $\beta$ in the following auxiliary regressions

$$
\begin{aligned}
(\pi-\hat{\pi}^{T}) &= \alpha(\hat{\pi}^{FHM}-\hat{\pi}^{T})+\hat{\vartheta} \\
(\pi-\hat{\pi}^{FHM}) &= \beta(\hat{\pi}^{T}-\hat{\pi}^{FHM})+\hat{\vartheta}'
\end{aligned}
\qquad [20]
$$

where $\hat{\pi}^{T}$ and $\hat{\pi}^{FHM}$ stand for the inflation forecasts produced on the basis of equations [17] and [19], respectively. These estimates provide for an indicator of the need to pool forecasts from both models. In the Forecast-Model Encompassing Test, instead, forecasts under FHM model are produced using only lagged information, i.e. using the marginal model $\pi_t \,/\, \pi_{t-1},\ \Delta w^*_{t-1},\ \Delta w^*_{t-2}, \Delta pm_{t-1},\ \Delta pm_{t-2},\ (p\text{-}w^*\text{-}\theta)_{t-1}$ rather than the conditional FHM model itself.

All statistics in the above tests are calculated on the basis of one-step ahead forecasts produced over the sample 1992:Q1-1997:Q4. We initially focus on one-quarter-ahead forecasts since this allows us to abstract from the lack of strong exogeneity required to produce dynamic forecasts at horizons longer than $h = 1$ in our single-equation framework. However, this exercise provides only a limited basis for evaluating forecasting performance, since relative rankings may not hold when longer horizons are considered.

**Figure 4** shows the resulting forecasts ($h = 1$) for GDP inflation (top panel), along with forecasts obtained at four (middle panel) and eight (bottom panel) quarters. **Figures 5a** and **5b** show the SET statistic recursively computed for T and FHM, respectively.

From a cursory look at the graphs, it is clear that while the FHM model (dotted line) fares comparatively well at the one-quarter horizon, its performance appears to deteriorate as the forecast horizon is extended. Conversely, T (solid thin line) appears to outperform FHM at the four-quarter horizon, in line with our previous findings concerning the location of the peak of the correlation between the real money gap and future inflation. At the eight-quarter horizon, none of the models performs well.

From a quantitative point of view, both equations produce unbiased forecasts on average. FHM in particular appears to anticipate inflation developments somewhat more precisely. The Root Mean Square Forecast Error turns out to be 0.77% for T and 0.60% for FHM. The results of our forecast encompassing tests, as shown in the table below, show that the SET tests cannot reject at standard confidence levels that both models T and FHM are valid simplifications from the minimal nesting model. This inference appears to be stable when the test statistics are computed recursively over the sample 1992:1 to 1997:4. Apparently, each model incorporates –though partially– some information that turns out to be relevant to explaining inflation developments in the euro area.

| Test | $H_0$: T | $H_0$: FHM |
|---|---|---|
| **SET Mizon-Richards** | $\chi^2(7) = 11.61(.110)$ | $\chi^2(4) = 2.79\ (.590)$ |
| **Forecast-Differential Encompassing** | F(1, 23) = 15.29** (.001)<br>$\alpha = .783\ (.188)$ | F(1, 23) = 1.170 (.291)<br>$\beta = .217\ (.188)$ |
| **Forecast-Model Encompassing** | F(6, 18) = 4.733** (.005) | F(4, 20) = .588 (.675) |

We finally test for out-of-sample forecast encompassing at longer time horizons, using only lagged information. More precisely, the two equations are re-estimated recursively over the sample 1992:Q1-1997:Q4 to produce 24 observations of out-of-sample forecasts at horizons ranging from $h=1$ to $h=8$ quarters[26]:

[26] In this exercise, the estimates of the different gap terms in [21] and [22] (and hence the long-run parameters) are, however, based on end-of-sample information. It should be noted that this way of proceeding ends up with downplaying the uncertainty surrounding inflation forecasts. Relevant to this latter point, see Orphanides and van Norden (1999) in relation to the measurement of the output gap.

$$(\pi_{t+h-1}-\pi_{t-1}) = \alpha_1^h(\pi_{t-1}-\hat{\pi}_t)+\alpha_2^h(\tilde{m}-\tilde{m}^*)_{t-1}+\alpha_3^h(y-y^*)_{t-1}+ \\ +\alpha_4^h[(R^s_{t-1}-\hat{\pi}_t)-(R^s-\pi)^*]+\hat{\varepsilon}_{1t+h-1} \tag{21}$$

$$(\pi_{t+h-1}-\pi_{t-1}) = \beta_0^h+\beta_1^h\pi_{t-1}+\sum_{i=1}^{2}\beta_{2i}^h\Delta w^*_{t-i}+ \\ +\sum_{i=1}^{2}\beta_{3i}^h\Delta pm_{t-i}+\beta_4^h(p-w^*-\theta)+\hat{\varepsilon}_{2t+h-1} \tag{22}$$

**Table 5** below shows Root Mean Square Forecast Errors at the different forecast horizons along with additional tests for forecast encompassing[27]. When compared to the equation standard errors reported above, the results for $h=1$ imply that both models fare particularly well over the sample considered (1992:1-1997:4), certainly because of the substantive decrease in inflation variability during that period. As regards forecasting encompassing, the table shows estimates of $\alpha$ (with standard errors) as well as $p$-values for the null hypotheses $H_0$: $\alpha=0$ and $H_1$: $\alpha=1$ in the auxiliary regressions:

$$\varepsilon_{1t}^h = \alpha(\varepsilon_{1t}^h-\varepsilon_{2t}^h)+\varsigma_t \tag{23}$$

where $\varepsilon_1^h$ and $\varepsilon_2^h$ ($h = 1,\ldots,8$) stand, respectively, for the forecast errors obtained from [21] and [22]. Rejection of one of the nulls would provide evidence in favour of the hypothesis that the rival model contains information helpful in explaining forecast errors from the own model, i.e. [21] under $H_0$ or [22] under $H_1$. Should this happen for both models, it would constitute evidence of some essential mis-specification in both models (Ericsson, 1992). This situation would call for forecast pooling, i.e., better overall inflation forecasts would be generated combining the forecasts obtained from both models.

The results in the table point out that the out-of-sample forecast record of our estimated model is comparatively better for all forecast horizons $h\geq 3$, though for $h\geq 6$ performance is rather poor for both models. Interestingly enough, however, the encompassing hypotheses are rejected at all horizons other than $h=1$. We interpret this outcome as a further indication that, at those horizons, pooling information from both models constitutes an improvement relative to forecasts generated from any of the models.

[27] See Chong and Hendry (1986) and Harvey et al. (1998).

| | | Model [21] | | Model [22] | |
|---|---|---|---|---|---|
| *h* | *α* | ***RMSE*** | *$H_0$: α=0* | ***RMSE*** | *$H_1$: α=1* |
| *1* | .830 (.348) | 0.653 | .0171** | 0.585 | .6251 |
| *2* | .575 (.176) | 0.656 | .0011** | 0.623 | .0156** |
| *3* | .241 (.053) | 0.638 | .0000** | 0.834 | .0000** |
| *4* | .178 (.098) | 0.614 | .0701* | 0.992 | .0000** |
| *5* | .335 (.121) | 0.827 | .0058** | 1.033 | .0000** |
| *6* | .454 (.094) | 1.094 | .0000** | 1.148 | .0000** |
| *7* | .447 (.157) | 1.145 | .0043** | 1.208 | .0004** |
| *8* | .472 (.198) | 1.259 | .0173** | 1.293 | .0078** |

**Table 5.** Root Mean Square Forecast Errors, alternative horizons; tests for forecast encompassing

## 6. Concluding Remarks

This chapter provided some preliminary empirical evidence on the information content of M3 broad money for future inflation in the euro area. Of course, serious measurement problems associated with these and other variables cannot be denied, but our attempt aimed at providing some evidence on the reality of day-to-day central banking, where information on the level of economic activity is generally incomplete.

First, we found little empirical support for rejecting at standard confidence levels Granger non-causality of M3 on prices. This conclusion is found stable throughout the sample and robust to a number of robustness checks.

Second, we investigated the leading indicator properties of broad money M3 by looking at a *Pstar*-type model in which information about the cyclical state of the economy and a measure of authorities' inflation target feed back onto the generation of inflation forecasts. In particular, our results confirm that a significant positive association exists in the euro area between the real money gap and future inflation up to five-six quarters ahead. Similar predictive ability is displayed by the output gap, although this finding, contrary to the previous one, does not prove to be robust to the use of alternative measures for the two unobservables. Finally, the real interest rate and (to a lesser extent) the term spread, appear to contain information that can be used to forecast inflation developments unexplained by recent extensions of baseline *Pstar* models for inflation.

One possible, though not exclusive, implication of the latter point is that the treatment of inflation expectations, and the measurement of the monetary policy authorities implicit inflation objective, are the key issue when it comes to providing structural explanations for the observed decline in area-wide inflation. Besides the predictive content of monetary indicators, the state of inflation expectations and the credibility of the policy stance appeared to be critical features of the inflation record in the period leading to, and immediately following, the start of EMU.

Third, we compared the forecasting ability of the model developed in the previous sections to that of an alternative, non-monetary model for euro-area GDP inflation by Fagan, Henry and Mestre (2000). The evidence we collected points to our modified *Pstar* model outperforming forecasts produced with the rival model at horizons $h \geq 3$, and being in turn outperformed at shorter horizons. However, our aim was not to compare alternative explanations of euro-area inflation. We rather wanted to perform a comparison between the forecasting ability of non-nested explanations of euro-area inflation in predicting price developments. Overall, each model appears to have some strengths of its own: both of them incorporate some information that is relevant to explain GDP inflation. However, taken individually, none of the models seems to be able to provide a complete account of inflation developments in the euro area. This clearly suggests that information derived from both monetary and non-monetary models (and indicators) should be used to generate reliable forecasts for euro-area inflation.

**Data Appendix**

The sample spans the years 1980Q1-1998Q4, and data come from European Central Bank Area-Wide Database. Series are seasonally adjusted aggregates from national sources, and in particular:

- GDP and GDP deflator are EUROSTAT and ECB aggregate data from national sources using fixed 1995 GDP weights at PPP rates

- Short- and long-term interest rates are BIS weighted averages of national rates using fixed 1995 GDP weights at PPP rates

- M3:
    - as from January 1999, ECB-calculated holdings of currency in circulation plus liabilities issued by MFIs and some central govt. institutions (overnight deposits, deposits with agreed maturity up to 2 years, deposits redeemable at notice up to 3 months, repos, money market fund shares, money market paper and debt securities with maturity up to 2 years); <u>series available from 1997Q3 onwards</u>
    - <u>1980Q1-1997Q2</u>: ECB aggregation of historical national estimates of the same series, with national contributions aggregated by their (fixed) GDP weight at PPP rates and 1997Q3 as starting values (ECB, 1999b)
    - <u>includes MFIs cross-border positions</u>
    - quarterly averages of monthly data

- Inflation is the annualised quarterly change in the GDP deflator, $\frac{\pi_t}{4} = p_t - p_{t-1}$
- Real money growth is the quarterly change in the real money stock, $\Delta\tilde{m}_t = \Delta\left(m_t - p_t\right) = \tilde{m}_t - \tilde{m}_{t-1}$

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| | **Granger non-Causality Tests** | | | **Wald tests for** | |
|---|---|---|---|---|---|
| | **M1:( m,p,y,s,l)** | | **M2:( m-p,y,s,l,p)** | **$H_0$: M2 vs $H_a$: M1** | |
| | I(2) model $d_{max}$=2 (1) | I(1) model $d_{max}$=1 (2) | I(1) model $d_{max}$=1 (3) | I(2) model $d_{max}$=2 (4) | I(1) model $d_{max}$=1 (5) |
| k=1 | 0.110 (0.740) | 0.970 (0.325) | 0.384 (0.535) | 9.109 (0.105) | 19.490 (0.109) |
| **k=2** | 0.579 (0.749) | 0.793 (0.673) | 0.958 (0.619) | 13.255 (0.151) | 9.837 (0.364) |
| k=3 | 1.102 (0.777) | 0.633 (0.889) | 1.066 (0.785) | 19.607 (0.106) | 20.599 (0.245) |

**Table 1** – Toda and Yamamoto (1995) tests for Granger non-causality of M3 on prices in the VAR(k). Standard errors in parentheses.

| | Tests for Granger non-Causality | | | | Wald tests for | |
|---|---|---|---|---|---|---|
| | Toda and Yamamoto | | | Toda and Phillips | $H_0$: M2 vs $H_a$: M1 | |
| | I(2) model $d_{max}=2$ | I(1) model $d_{max}=1$ | I(1) model $d_{max}=1$ | | I(2) model $d_{max}=2$ | I(1) model $d_{max}=1$ |
| | **M1:(m,p)** | | **M2:(m-p,$\pi$)** | **M2:(m-p,$\pi$)** r=0 | | |
| k=1 | 0.004 (0.950) | 0.134 (0.714) | 0.425 (0.515) | H*=2.045 (.153) | 8.189** (0.017) | 23.932*** (0.000) |
| k=2 | 0.814 (0.666) | 1.410 (0.494) | 2.308 (0.316) | | 5.105* (0.078) | 6.737** (0.034) |
| k=3 | 0.257 (0.968) | 2.273 (0.518) | 1.819 (0.611) | | 2.0599 (0.357) | 0.059 (0.970) |
| | **M1:(m,p,y)** | | **M2:(m-p,$\pi$,y)** | **M2:(m-p,$\pi$,y)** r=1 | | |
| k=1 | 0.097 (0.755) | 0.006 (0.938) | 0.614 (0.434) | $H_1$*=15.71 (.000)<br>$H_2$*=2.22 (.136) | 4.721 (0.193) | 11.357** (0.010) |
| k=2 | 0.518 (0.772) | 0.950 (0.622) | 1.033 (0.597) | $H_T$*=.345 (.557) | 6.061 (0.109) | 15.633*** (0.001) |
| k=3 | 3.387 (0.336) | 1.159 (0.763) | 0.827 (0.843) | H*=3.969 (.137) | 2.150 (0.542) | 1.690 (0.639) |
| | **M1:(m,p,y,s)** | | **M2:(m-p,$\pi$,y,s)** | **M2:(m-p,$\pi$,y,s)** r=2 | | |
| k=1 | 0.970 (0.325) | 2.470 (0.116) | 1.337 (0.248) | $H_1$*=17.71 (.000)<br>$H_2$*=5.20 (.074) | 5.209 (0.267) | 5.158 (0.272) |
| k=2 | 0.846 (0.655) | 1.047 (0.592) | 0.420 (0.811) | $H_T$*=.013 (.909) | 2.104 (0.717) | 2.699 (0.609) |
| k=3 | 2.104 (0.551) | 1.046 (0.790) | 0.381 (0.944) | H*=1.556 (.459) | 2.970 (0.563) | 1.935 (0.748) |
| | **M1:(m,p,y,l-s)** | | **M2:(m-p,$\pi$,y,l-s)** | **M2:(m-p,$\pi$,y,l-s)** r=2 | | |
| k=1 | 0.192 (0.661) | 0.466 (0.495) | 0.309 (0.578) | $H_1$*=22.72 (.000)<br>$H_2$*=2.59 (.274) | 4.601 (0.331) | 4.215 (0.378) |
| k=2 | 0.243 (0.886) | 0.335 (0.846) | 0.445 (0.801) | $H_T$*=.067 (.796) | 4.726 (0.317) | 11.131** (0.025) |
| k=3 | 0.312 (0.956) | 0.810 (0.847) | 0.527 (0.913) | H*=2.74 (.253) | 1.819 (0.769) | 1.593 (0.810) |

**Table 2** – Toda and Yamamoto(1995) and Toda and Phillips (1994) tests for Granger non-causality of M3 on prices in the VAR(k). Standard erors in parentheses.

| Horizon → Parameter↓ | *h=1* | *h=2* | *h=3* | *h=4* | *h=5* | *h=6* | *h=7* | *h=8* | *h=9* | *h=10* | *h=11* | *h=12* |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| $\hat{\alpha}_1^h$ | .217** | -.012 | -.124 | -.053 | -.147 | -.159 | -.114 | .127 | -.146 | -.172 | -.191 | -.035 |
| $\hat{\alpha}_2^h$ | .196* | .283* | .355** | .341** | .316** | .205* | .028 | -.119 | -.076 | -.076 | -.087 | -.206 |
| $\hat{\alpha}_3^h$ | .262** | .412** | .533** | .588** | .652** | .524** | .234 | .096 | .178 | .240 | .211 | .161 |
| $\hat{\alpha}_4^h$ | .249* | .255* | .196** | .097 | .113 | .146 | .211 | .141 | .212** | .149 | .096 | -.003 |
| $\overline{R}^2$ | .52 | .45 | .41 | .33 | .32 | .21 | .11 | .11 | .09 | .06 | .03 | .07 |

**Table 3** – Forecasting equations for inflation, estimated coefficients at different forecasting horizons. ** and * indicate significance at 5 and 10% confidence levels

| | $\hat{\pi}$ = deterministic exponential trend | | | | $y^*$ = HP(1600) | | | |
|---|---|---|---|---|---|---|---|---|
| | $y^*$= AWM | $y^*$= HP(1600) | $y^*$= HP(5000) | $y^*$=HP(15000) | $\hat{\pi}$ = determ. exp. trend | $\hat{\pi}$ = HP(1600) | $\hat{\pi}$ = HP(5000) | $\hat{\pi}$ = HP(15000) |
| $\hat{\theta}_0$ | **-.774**** | **-.812**** | **-.811**** | **-.809**** | **-.812**** | **-.831**** | **-.827**** | **-.814**** |
| | (.110) | (.108) | (.107) | (.107) | (.108) | (.114) | (.112) | (.110) |
| $\hat{\theta}_1$ | **.148**** | **.210**** | **.197**** | **.170**** | **.210**** | **.064** | **.109*** | **.175**** |
| | (.068) | (.076) | (.070) | (.065) | (.076) | (.056) | (.062) | (.073) |
| $\hat{\theta}_2$ | **.065** | **.264*** | **.154** | **.095** | **.264*** | **.195** | **.245*** | **.291**** |
| | (.086) | (.134) | (.096) | (.083) | (.134) | (.132) | (.136) | (.141) |
| $\hat{\theta}_3$ | **.340**** | **.211** | **.291**** | **.320**** | **.211** | **.226** | **.210** | **.188** |
| | (.141) | (.133) | (.132) | (.134) | (.133) | (.136) | (.136) | (.138) |
| $\hat{\theta}_4$ | **.446**** | **.348**** | **.362**** | **.366**** | **.348**** | **.271**** | **.282**** | **.281**** |
| | (.092) | (.103) | (.100) | (.101) | (.103) | (.116) | (.113) | (.113) |
| | $\sigma$=.80% | $\sigma$=.79% | $\sigma$=.78% | $\sigma$=.78% | $\sigma$=.79% | $\sigma$=.80% | $\sigma$=.80% | $\sigma$=.80% |
| | $R^2$=.45 | $R^2$=.47 | $R^2$=.48 | $R^2$=.48 | $R^2$=.47 | $R^2$=.45 | $R^2$=.46 | $R^2$=.46 |

**Table 4** - Forecasting equation for inflation, estimated coefficients at different forecasting horizons. ** and * indicate significance at 5 and 10% confidence levels. AWM stands for the measure of potential output included in the ECB Area-Wide Model. HP stands for Hodrick-Prescott filter, for various smoothing parameters $\lambda$.

Figure 1a.

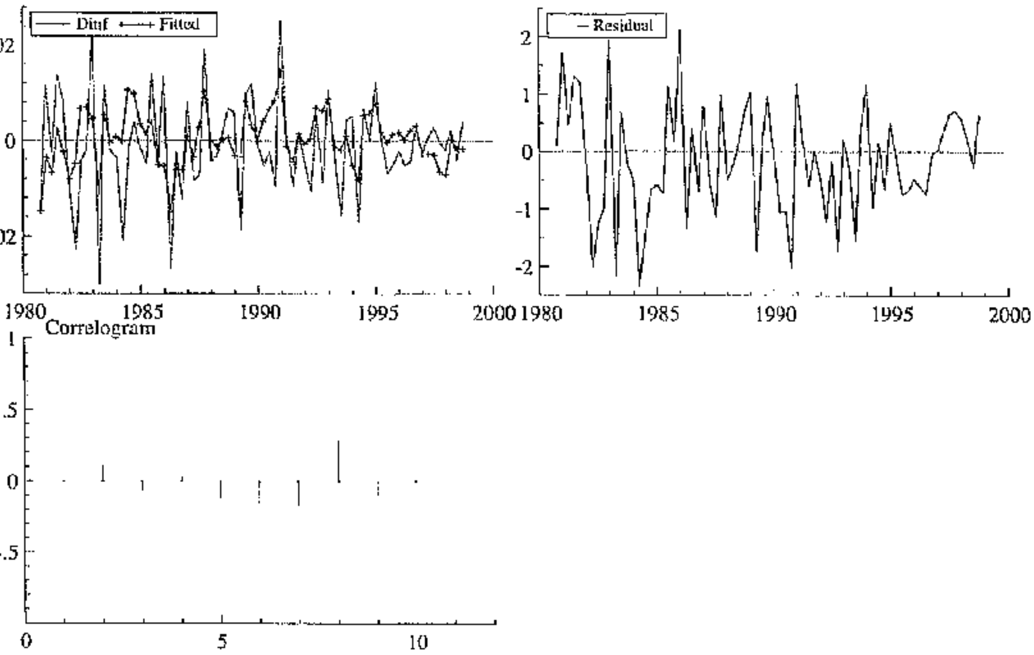



Figure 1b.

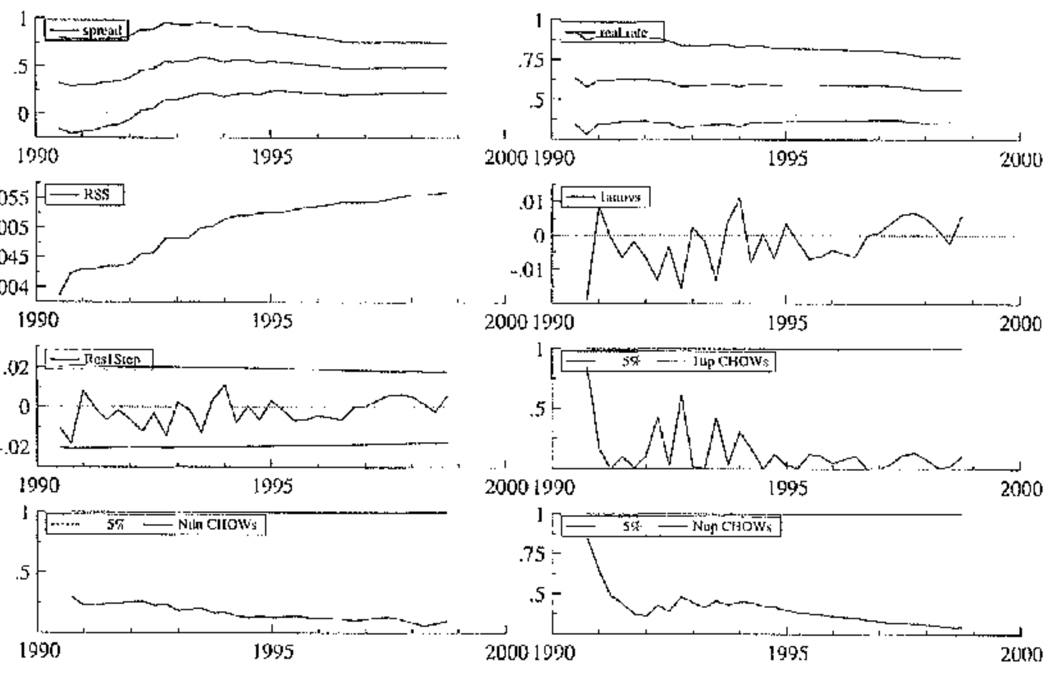


Figure 2a

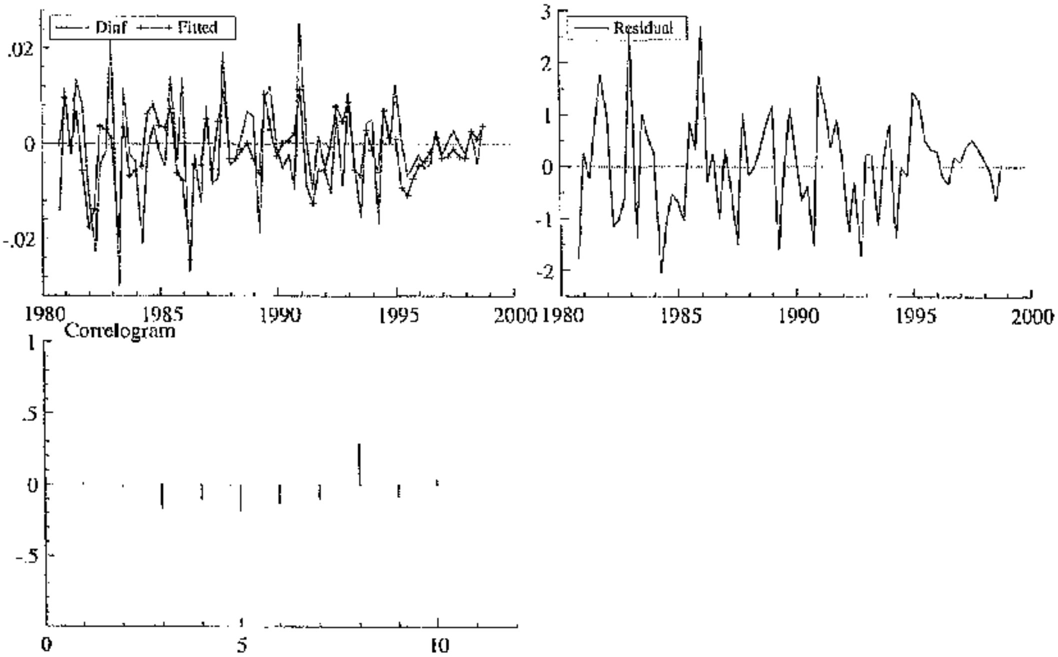
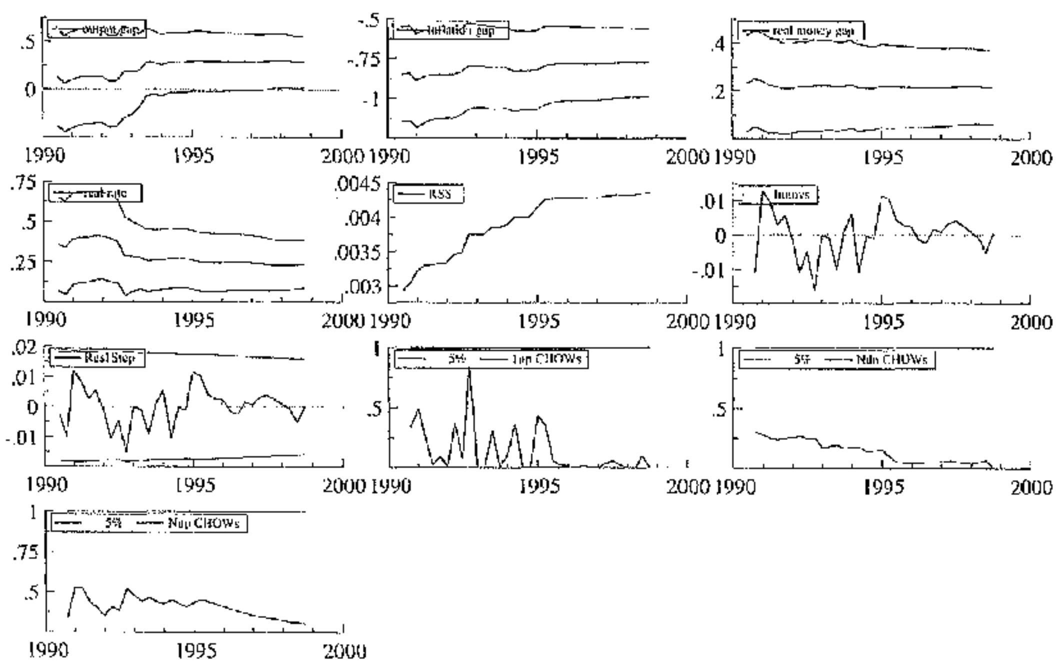



Figure 2b



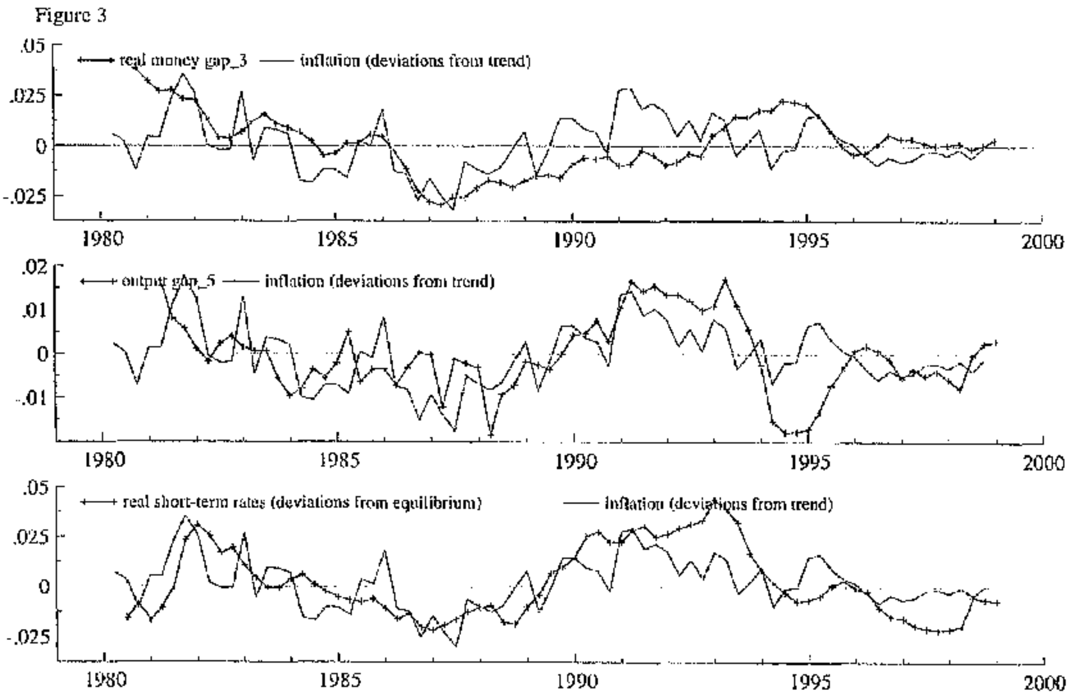



**Figure 3.** Lagged real money gap, output gap and real short-term interest rates versus inflation

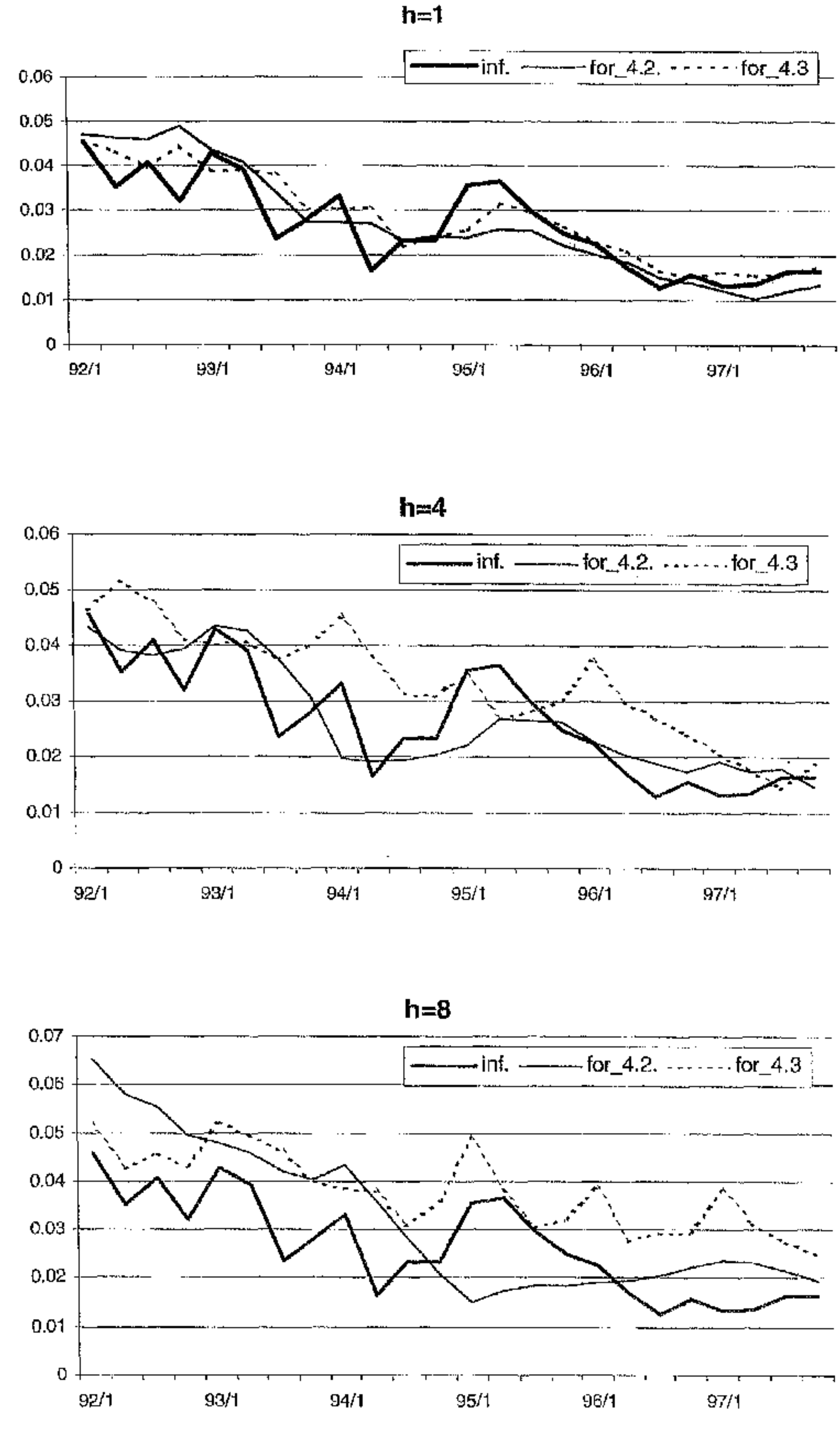



**Figure 4** – Inflation (bold line), and inflation forecasts from T (solid thin line) and FHM (dotted line) at various horizons h.

Figure 5a. SET: recursive computation (Ho: 3.7)

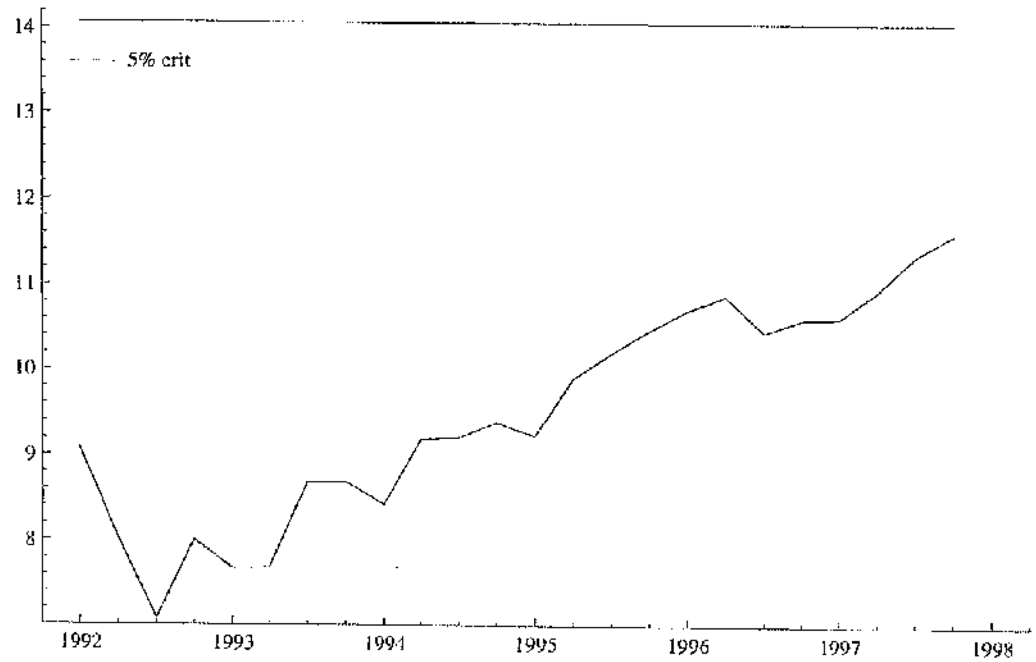



Figure. 5b SET: recursive computation (Ho: 4.1)

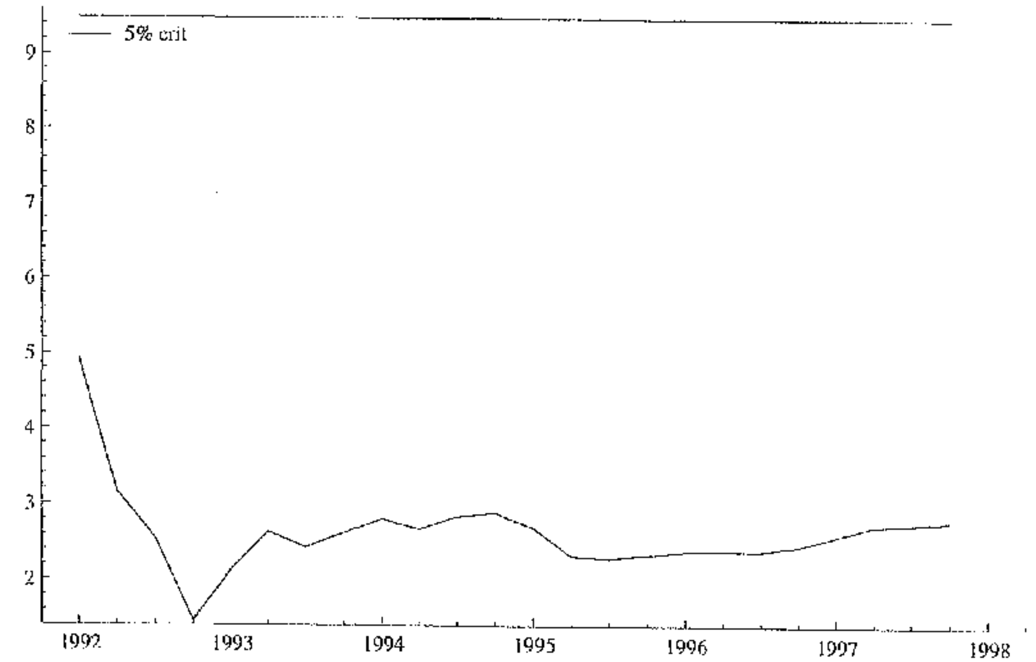



**Figure 5** – SET, recursive computation.

# Time-Varying-VAR Perspectives on Real Exchange Rates, Productivity Levels and Government Spending

# Chapter 4

> *"...The starting point of the approach is that in specifying econometric models one attempts to filter as much information as possible from the data. However, instead of relying on classical hypothesis testing or economic theory to decide whether a particular variable or lag should enter the autoregression, these authors use a symmetric "atheoretical" prior on all variables to trade off overparameterisation with oversimplification. The reason for taking an alternative route to the specification problem is that there is a very low signal-to-noise ratio in economic data and economic theory leaves a great deal of uncertainty concerning which economic structures are useful for inference and forecasting. Because highly parameterised unrestricted VAR models include many variables in each equation, the extraction filter is too wide and noise obscures the relatively weak signal present in the data: the prior acts as an orientable antenna which, when appropriately directed, may clarify the signal..."*
>
> Canova (1995), pp.84-85

## 1. Introduction

Over the past two decades, developments in econometric theory have made possible to achieve a better understanding of the reasons why substantial deviations of the exchange rate from the Purchasing Power Parity (PPP) can take place. In particular, the analysis of this phenomenon has greatly benefited from the emergence of concepts like stationarity and cointegration, and from their application to the study of exchange rates' trending properties.

Although a more decisive response on this issue will need to resort to further advances in econometric theory, the baseline findings of this voluminous body of empirical literature can be summarised as follows. First, PPP appears to hold, but only in the very long run: both in its absolute and relative versions, PPP fails to hold continuously (Clarida and Gali, 1994; MacDonald, 1996; Engel, 1996). Second, the observed departures of the exchange rate from PPP are more persistent than traditional, flexible-price models of the exchange rate would predict.

As a direct consequence of such findings, the focus of the theoretical literature on exchange rates has recently shifted towards the development of intertemporal sticky-price models, in an effort to take more adequate consideration of the dynamic adjustment followed by saving and investment (Obstfeld and Rogoff, 1997).

This chapter aims at providing some empirical evidence as to the *origins* of the observed deviations of the exchange rate from PPP. The question of whether PPP holds, and the related one about the mean-reverting properties of the nominal exchange rate, yet deserve extensive empirical work, but our focus is different. In what follows we wish to shed some light on the *nature* of the movements of the real exchange rate.

Thus far, the existing empirical evidence has provided some limited support to the idea that sustained divergences in government spending and sectoral productivity patterns might be at the root of persistent fluctuations of bilateral real exchange rates. Moreover, the persistence of deviations of the nominal exchange rate from PPP appears to critically depend on the exchange rate regime in place. That is, during periods of floating exchange rates, persistent fluctuations of the real exchange rate seem to be more common than under fixed exchange rates, thus producing the largest observed departures from PPP[1]. Finally, market imperfections -in the form of pricing-to-market behaviour or similar market segmentation practices- may prevent perfect goods arbitrage, and make the dynamics of relative prices diverging from that of nominal exchange rates.

Against this background, we then seek to attribute the observed movements of the bilateral real exchange rates between the USA, UK, and Italy, to some of the causes above summarised. That is, we try a tentative attribution of the shocks to the real exchange rates to existing divergences in the fiscal stance[2], differential productivity levels, and significant differences in money market conditions.

There are two main novelties in our study. First, following some recent advances in Bayesian approaches to the estimation of Vector Autoregressions, we employ a time-varying methodology. We apply such techniques to the estimation of an unrestricted VAR comprising the bilateral real exchange rate, a measure of productivity differentials, an indicator of the relative fiscal stance, and the ratio between real *ex-post* interest rates in the two countries. We then decompose the total residual variance of each VAR equation into stochastic contributions attributable to innovations in each endogenous variable. By applying a Kalman filter technique to the system's estimated parameters and variances, we decompose the total variance of the real exchange rate equation into contributions that are allowed to change over time. The particular state-space representation we adopt for our

[1] Figure 1 displays series for the three bilateral real exchange rates we study. From it, it is relatively apparent that the major fluctuations of the rates are concentrated during the free-floating intervals.
[2] Here we do not distinguish between the fiscal balance *per se* and the unexpected fiscal stance shocks as the emerging literature on fiscal policy transmission seeks to identify (Fatas and Mihov, 1998; Blanchard and Perotti, 1999; Giavazzi

models enables us to impose the least restrictive assumptions on the dynamic structure of the system; the VAR coefficients are simply assumed to be stationary white noise processes. This way, our model is able to pick up the changing influences of the monetary regime, of productivity shocks, and the fiscal and monetary stance, on the real exchange rate, over time. We further avoid to impose particularly stringent identification structures, by adopting the Generalised Impulse Response approach of Koop *et al.* (1996) and Pesaran and Shin (1998) to examine the dynamic response of the real exchange rate to shocks in each of the remaining variables of the system.

The second novelty of this study is represented by the data we use. We apply our approach to a long historical set of annual observations spanning the last 130 years and obtained from a variety of sources. The bilateral real exchange rates between the USA, UK, and Italy are analysed across all international monetary regimes (classical Gold Standard, Bretton Woods, the post-1973 floating and the EMS) these countries were involved in during the sample period.

There is a clear policy-motivated aim in the kind of analysis we undertake. Recently, monetary policymaking has been unified in Europe, with the introduction of the euro and the establishment of the ESCB. Moreover, recurrent proposals of dollarisation are put forward for some developing and transition economies. Should real shocks emerge as the driving force of real exchange rate fluctuations, this would constitute a relatively negative message for such attempts, and would instead provide a strong case in favour of flexible exchange rate regimes

The main results we obtain can be summarised as follows. First, we find very little evidence in favour of shocks to relative productivity levels as having persistent appreciating effects on real exchange rates. Such result, though not valid across all the sub-sample estimates we obtain, appears to be overall robust. Second, we find some stronger support for shocks to the relative fiscal position of a country as triggering substantial real exchange rate fluctuations. Third, estimates conducted within single exchange rate regimes suggest that the response of real exchange rates to shocks in the other variables appears to critically depend not just on the international monetary arrangements historically in place, but also on demographic and technological factors. As can be seen, the above mentioned dilemma between fixed and flexible exchange rate regimes is not fully solved by the evidence we uncover, and calls for further investigation.

et al. , 2000). The distinction is very important but, for sake of simplicity, here we refer to the fiscal stance generically, The structure of this chapter is as follows. **Section 2** recalls a simple analytical framework commonly used for standard testing of the PPP hypothesis, and then turns to the existing empirical evidence on real exchange rates dynamics, briefly summarising the main findings so far. **Section 3** formalises the small VAR model of the real exchange rate adopted for our following empirical investigation, and discusses various identification issues related with the problem at hand. Then, our time-varying framework is introduced. **Section 4** describes in detail the econometric techniques we use, while **Section 5** illustrates our main findings. Finally, **Section 6** contains some concluding remarks and directions for further empirical research.

## 2. Modelling Real Exchange Rates in the Long Run

### *2.1 A Simple Theoretical Framework*

The framework used by the recent wave of tests on the PPP hypothesis rests on a simple specification of production and consumption in an intertemporal context. Before we start discussing the existing empirical evidence on this issue, we introduce a simple theoretical model that encompasses several of the features the literature has recently highlighted. In what follows, we illustrate a baseline framework first introduced by Rogoff (1992), and slightly modified for estimation purposes by Chinn and Johnston (1996).

The economy consists of two sectors, in which output is determined by Cobb-Douglas-type production functions:

$$\begin{aligned} Y_t^T &= A_t^T \left(L_t^T\right)^{\phi T} \left(K_t^T\right)^{1-\phi T} \\ Y_t^N &= A_t^N \left(L_t^N\right)^{\phi N} \left(K_t^N\right)^{1-\phi N}, \end{aligned} \qquad [1]$$

where $Y^T$ and $Y^N$ are output of the traded and nontraded sectors, respectively, $K$ and $L$ are capital and labour inputs, and $A$ is a stochastic productivity shock. In this as in many other similar models, the existence of some adjustment cost prevents, in the short-run, full factor mobility.

The general price level is a geometric average of the traded and nontraded goods price indices (all small-cap variables are in logs):

deferring the question to some future research.

$$p_t = \alpha p_t^N + (1-\alpha) p_t^T \qquad [2]$$

The representative agent in the economy maximises a utility function expressed in terms of consumption of the two goods, $C^T$ and $C^N$:

$$U_t = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left\{ \frac{\left[ \left(C_s^N\right)^{\theta} \left(C_s^T\right)^{1-\theta} \right]^{1-\lambda}}{1-\lambda} \right\}, \qquad [3]$$

where $\beta$ is the subjective discount rate, and $\lambda$ is the inverse of the elasticity of intertemporal substitution. Elementary budget constraints for private and government consumption hold, so that in each period total consumption in each sector equals output in each sector. Consequently, the price of nontraded goods relative to traded ones, $\tilde{P}_t \equiv \frac{P_t^N}{P_t^T}$, is a function of relative consumption of the two goods:

$$\tilde{P}_t = \frac{\theta C_t^T}{(1-\theta) C_t^N} \qquad [4]$$

Agents smooth expected marginal utility over time, so that the first order condition from the optimisation problem can be approximated by an expression in which we also assume that the *A*s have constant variance:

$$E_t\left(c_{t+1}^T - c_t^T\right) = \frac{\theta(1-\lambda)}{\lambda + \theta(1-\lambda)} E_t\left(c_{t+1}^N - c_t^N\right) \qquad [5]$$

Rogoff (1992) shows that, by combining [3] into [4] and rearranging, relative price inflation becomes proportional to the growth of relative demand for the two goods:

$$\tilde{p}_{t+1} - \tilde{p}_t = \left(c_{t+1}^T - c_{t+1}^N\right) - \left(c_t^T - c_t^N\right) \qquad [6]$$

We now further assume that government spending falls entirely on nontraded goods and that it behaves according to a random walk. Clearly, the former assumption, along with the assumed existence of some adjustment cost affecting the reallocation of production factors between sectors, is the key to determine persistent effects of demand-side shocks on the real exchange rate. Alternatively, one might assume the existence of market power, or the breakdown of PPP for traded goods, to obtain the same effects[3].

If technology shocks are normally distributed, that is

$$\begin{aligned} a_{t+1}^N &= a_t^N + \varepsilon_t^N \\ a_{t+1}^T &= a_t^T + \varepsilon_t^T \end{aligned}, \qquad [7]$$

one can amend equation [6] to obtain an expression in which relative price inflation is also function of the ratio of nontraded output to private demand for it. If $f_N$ is such ratio, [6] can be re-written as follows:

$$\tilde{p}_{t+1} - \tilde{p}_t = \left(a_{t+1}^T - a_t^T\right) - f_N\left(a_{t+1}^N - a_t^N\right) + \left(f_N - 1\right)\left(g_{t+1} - g_t\right), \qquad [8]$$

where $g_t$ is (the log of) government spending. After recursive backward substitution, we obtain

$$\tilde{p}_{t+1} = a_{t+1}^T - f_N a_{t+1}^N + \left(f_N - 1\right) g_{t+1} + k_0, \qquad [9]$$

where $k_0$ are the initial relative price conditions.

We now introduce the foreign country in our model. We start by defining the real exchange rate ($q$) as the nominal exchange rate ($e$) multiplied by the ratio between aggregate price indices: $q_t = e_t + p_t^* - p_t$. We further assume that the foreign country is identical to the home one -which means that for the moment we assume the two countries as having the same weight $\alpha$ in their price indices- and that PPP holds for traded goods. Through manipulation of the expression for the real exchange rate, we obtain:

[3] De Gregorio, Giovannini and Wolf (1994) pursue this lines of reasoning quite effectively in their empirical investigation, surveyed below

$$q_t = \alpha\left(e_t + p_t^{N^*} - p_t^N\right), \tag{10}$$

where now the real exchange rate is a function of the relative price of nontraded goods.

Using the previous equation, we can express [9] in terms of departures of home relative prices from foreign relative prices:

$$\begin{aligned}\tilde{p}_{t+1} - \tilde{p}^*_{t+1} &= \left(p^N_{t+1} - p^T_{t+1}\right) - \left(e_t + p^{N^*}_{t+1} - e_t - p^{T^*}_{t+1}\right) = \\ &= \widehat{a}^T_{t+1} - f_N \widehat{a}^N_{t+1} + \left(f_N - 1\right)\widehat{g}_{t+1} + \widehat{p}_0\end{aligned} \tag{11}$$

In the above expression, starred variables refer to the foreign country, while the circumflex indicates a variable calculated in terms of relative differences (home minus foreign). In [11], the LHS is simply the ratio between relative price levels in the two countries. The fact that PPP holds for traded goods implies that:

$$\left(e_{t+1} + p^{N^*}_{t+1} - p^N_{t+1}\right) = \widehat{a}^T_{t+1} - f_N \widehat{a}^N_{t+1} + \left(f_N - 1\right)\widehat{g}_{t+1} + \widehat{p}_0, \tag{12}$$

and then, using [10]:

$$q_{t+1} = -\alpha[\widehat{a}^T_{t+1} - f_N \widehat{a}^N_{t+1} + \left(f_N - 1\right)\widehat{g}_{t+1} + \widehat{p}_0] \tag{13}$$

The above equation illustrates quite well a number of features of the long-run relationship between the real exchange rate, the relative sectoral productivity levels, and government spending as a fraction of GDP. This relationship, or very similar ones, has often been used to develop testable hypotheses about real exchange rate movements. For example, diverging trends in the relative productivity levels in the traded and nontraded goods sectors can be seen in [13] as triggering an appreciation of the real exchange rate. According to the well-known Balassa-Samuelson hypothesis (BSH, henceforth. See Balassa, 1964; Samuelson, 1964), if in one of the countries productivity tends to rise more in the traded goods sector than in the nontraded goods one, the relative price of nontraded goods will tend to rise. This because of, assuming perfect international capital mobility, productivity gains, and subsequent production cost savings, in the traded goods sector. It follows that the general CPI will rise more under the pressure of non-traded goods inflation. This will lead the price levels in economies that are experiencing productivity gains or are catching-up with more advanced countries to rise more quickly than in countries enjoying their steady-state levels of productivity growth. Real exchange rate appreciations on the part of faster-growing economies relative to steady-state ones should thus ensue.

It is useful to point out that this result holds thanks to a series of standard, but critical assumptions. For example, the above model explicitly assumes that capital is perfectly mobile internationally, that capital and labour are perfectly mobile across sectors, and that Ricardian equivalence holds (Obstfeld and Rogoff, 1997)[4].

Rogoff (1992) developed this line of reasoning slightly further, by noting that in [11] temporary shocks to relative productivity can have persistent or even permanent effects on the relative price of nontraded goods. This because internationally developed capital markets will help agents to smooth out their consumption of traded goods. The intratemporal relative price of nontraded goods will then be smoothed too.

The natural question about the effect of diverging productivity patterns on the price of nontraded goods is then the following: should we expect to detect a Balassa-Samuelson effect in the long run? The final answer is eminently empirical, and indeed the evidence reviewed below provides some clue about it. However, two stylised facts should be borne in mind. First, technological advances tend to spread very fast, and the catching-up of productivity levels in some previously lagging economies is one of the best-known historical facts of the 19$^{th}$ and 20$^{th}$ centuries. As lagging economies approach steady-state levels of productivity, the Balassa-Samuelson effect should gradually die away. Interestingly, the US and Italy were, for parts of our sample, experiencing catching-up processes vis-à-vis a more mature industrial country like the UK. In turn Britain has seen its productivity levels relative to the US and Italy undergoing substantial fluctuations in the past 130 years. More importantly, even if productivity levels can diverge across countries for prolonged periods, one expects capital and labour mobility pushing towards long-run convergence of income levels. It is then clear how using long-span, low frequency data, is the only way one has to draw some conclusions about the persistence of such effects in the long run.

[4] If Ricardian equivalence holds, a temporary tax has no effect on aggregate demand or the exchange rate. An additional implicit assumption is that taxation does not generate distortionary effects on private spending and labour supply.

However, Froot and Rogoff (1995) slightly generalise this model, obtaining a formulation in which the change in the relative price of traded goods is also a function of each sector's relative capital intensity -the $\phi$ parameters in our equations [1]. They thus point out that, even in the case of balanced productivity growth between the traded and non-traded sectors, if the production of non-traded goods is more labour-intensive, the relative price of nontraded goods will rise, triggering once again an appreciation of the real exchange rate.

As far as the effects of demand factors on the real exchange rate are concerned, from [13] we see that since government spending is assumed to be falling entirely on nontraded goods, the relative price of the latter, and the real exchange rate, are bound to appreciate. Clearly, this holds only as long as capital and labour are not perfectly mobile across sectors, so that the effects of any demand-side shock on the relative price of nontraded goods should be limited to the short run only. But this point, as well as the one related to the persistence of the effects of monetary shocks on the real exchange rate, appears another purely empirical question.

An additional practical point relates to the difficulty of devising productivity measures that are unrelated to the measures of fiscal stance, and ultimately, to the business cycle. As we will see, this may prevent, in principle, clear identification of the transmission of the shocks to the real exchange rate.

We now turn to analysing the empirical evidence about exchange rate deviations from PPP.

### *2.2 Productivity Differentials, Fiscal Shocks and Real Exchange Rate Movements*

What follows is a brief review of the most recent empirical studies on the effects of differences in productivity levels and shocks to the fiscal stance on real exchange rates. As surveys with similar aims are widely available[5], we will only focus on particular aspects of each contribution.

De Gregorio, Giovannini and Wolf (1994) represents one of the first extensive studies about the relative importance of demand and supply factors in determining real

Alesina and Perotti (1995) and Giavazzi, Jappelli and Pagano (2000) somehow elaborate on the empirical relevance of these effects.

exchange rate movements. They start by constructing some measures of total factor productivity growth -as opposed to the labour productivity measure used in previous analyses- for 14 countries, using the OECD intersectoral database. After providing an innovative classification of the sectors in traded and nontraded, the authors pool data for all the countries over the period 1971-1985. De Gregorio *et al.* (1994) then detect significant effects of productivity growth and government spending as a share of GNP on the relative price of nontraded goods. However, only the former remains significant when similar regressions are run on cross-sectional data averaged for each country –a way of solving the baseline model for the long-run equilibrium. That is, the productivity effect on real exchange rates is found to be persistent.

Using a similar data set, Asea and Mendoza (1994) find that the BSH helps explaining the observed changes in the relative price of nontraded goods, in a sample of 14 OECD countries between 1975 and 1985. However, such price changes do not appear to be relevant in explaining measured real exchange rate fluctuations over that sample.

Chinn (1997) and Chinn and Johnston (1996) assess whether the apparent lack of cointegration usually found between time series of real exchange rates, sectoral productivity levels and the fiscal balance[6], is replicated when panel cointegration techniques are applied. Their findings are that panel data tend to confirm such cointegrating relationship, and in particular that productivity shocks in the traded goods sector do trigger a real appreciation in the real exchange rate. At the same conclusion arrived similar panel cointegration tests by Canzoneri, Cumby and Diba (1996) on a different panel of OECD economies. In fact, their evidence is particularly supportive of the real exchange rate being cointegrated with sectoral labour productivity differentials. In turn, Alberola and Tyrväinen (1998) use standard cointegration analysis to verify the same hypothesis for a number of European countries. Their findings confirm the results in Chinn (1997)[7], though they reject the complete equalisation of wages across sectors that usually accompanies the baseline BSH.

Canzoneri, Cumby, Diba and Eudey (1998) look at the empirical evidence in favour of the BSH, for a number of EU countries. Their informal examination of data on relative prices of home and traded goods reveals two basic points. First, although overall

[5] Asea and Corden (1994) and Froot and Rogoff (1995) provide exhaustive reviews of this literature, more recently assessed by MacDonald (1999).
[6] Our equation [13] can be easily interpreted as such cointegrating relationship.
[7] Their econometric results find support for the BSH in Germany, Belgium and Spain. For France, Italy, Netherlands, Austria and Finland the BSH is however rejected.

inflation has converged among the EU Member States, there are still substantial differentials between non-traded goods inflation and traded goods inflation. In addition, differences are significant in home-good inflation rates between Germany and the other economies. In other words, all EU countries display significant nontraded goods inflation differentials. Secondly, productivity differentials with Germany seem to be the most important explanatory factor in generating inflation differentials amongst sectors.

Formal unit-root tests conducted by Canzoneri *et al.* (1998) on a panel of all EU countries confirm that trends in relative prices of nontraded goods are to be explained by trends in relative productivity across sectors. This validates the BSH for the countries and the sample at hand.

Against this background, however, two points are worth noting. As argued, *inter alia*, by Rogoff (1995, 1996), if the law of one price fails to hold within the traded goods sector, the aggregate price ratio between countries will not move in line with the nominal exchange rate. Consequently, deviations from PPP tend to exist even in the absence of diverging productivity patterns or fiscal policy shocks. In addition, monetary factors have not, so far, been considered in the analysis. It is instead true that they may play a major role in determining prolonged fluctuations in nominal, as well as in real exchange rates[8]. It might be more appropriate to think of the relationship between productivity differentials, the fiscal balance and the real exchange rate as not invariant, for example, to the exchange regime in place, or to the monetary stance more in general. We generally do expect monetary shocks to have only temporary effects on real exchange rates[9]. However, the monetary regime in which the country under investigation finds itself is likely to determine the transmission channels of productivity and fiscal balance shocks to the real exchange rate.

We then turn to assess the ways in which the empirical literature has faced these two issues.

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[8] Rogers (1999), as we will see, provides some evidence in favour of such a case.

[9] It is true, however, that every business cycle factor permanently affecting employment and output can have persistent effects on the real exchange rate too. This can happen through permanent changes in long-run productivity (Aghion and Saint-Paul, 1993), vintage effects in investment and hysteresis in the labour market (Darby, Ireland and Wren-Lewis,

## 2.3 *Market Imperfections and Monetary Shocks: Do They Matter?*

The failure of the law of one price is now a standard result in a number of empirical open-economy studies. Campa and Wolf (1998) argue that the mean-reversion properties of G-7 economies' real exchange rates can be attributed only to a limited extent to goods market arbitrage. Rogoff's (1996) conclusion is that international product markets are probably still segmented to such an extent to produce large trading frictions and nominal rigidities. The same explanations are used by Cecchetti, Nelson and Sonora (1998) to illustrate their findings, which prove the existence of a PPP puzzle for relative prices within the USA as well.

In line with the latter findings, Engel and Rogers (1998b) show that relative prices between two locations tend to be indefinitely more volatile than the simple geographical distance between markets can otherwise explain. The study stresses the importance of alternative factors in determining how close the integration of consumer markets is between different locations. Amongst these, the fraction of non-traded components -mainly local services- in any product; the eventual existence of barriers to shipments for the traded components; finally, the degree of monopolistic mark-up. The third of these factors, in particular, proxies for pricing-to-market behaviour on the part of firms and allows for the possibility that market segmentation may provide room for price discrimination. The basic findings of this research tend to show that relative price variability is well explained by a geographical-distance effect, and that the "tradability characteristics" of the good play no role in that. Another contribution by Engel and Rogers (1998a), testing the same model on a sample of 23 countries on different continents, finds that international relative price volatility is a function of both distance and exchange rate variability. On the other hand, Engel, Hendrickson and Rogers (1997)[10] reject the presence of a unit root in real exchange rates and show that the speed of convergence to PPP is different for intra-national, intra-continental and inter-continental real exchange rates.

Engel and Rogers (1998b) allow for the possibility that nominal exchange rate volatility might explain much of the cross-border price variability. If nominal prices are sticky and the nominal exchange rate is highly volatile, cross-country prices will be more volatile than within-country prices, simply because the former are expressed through a

1996), or simply the loss of skills normally associated with severe business cycle downturns. These points were raised by Muscatelli and Spinelli (1999).

[10] They use panel data from four countries on two continents.

nominal exchange rate measure. In fact, Bayoumi and MacDonald (1999), by applying panel unit root methods, detect this effect in the context of two monetary unions, Canada and the United States. More generally, this pattern re-emerges in the comparison of intra-continental to cross-continental relative price: the intra-continental prices display lower variances (Engel and Rogers, 1998a). In fact, the border effect appears solid even to this caveat, and the simple relative stickiness of national prices is not able to fully explain market segmentation.

Cecchetti, Nelson and Sonora (1998) reach the same conclusions. Their main objective is to test for general price level discrepancies across U.S. cities, and they find large and persistent divergences. For example, annual inflation rates measured over 10-year intervals are shown to differ by up to 1.6 percent. In the context of EMU, such differentials might involve substantial complications for ESCB's interest rate policy.

In surveying all the above analyses and those in the previous section, we have devoted little attention to the nature of the monetary regime in place. In fact, one of the stylised facts mentioned in the introduction is that, during periods of floating exchange rates, real exchange volatility appears to be higher. Indeed, using historical data on EMU Member States as well as non-EMU countries, Obstfeld (1998a, b) shows that, in low-to moderate-inflation contexts, shifts from floating to controlled nominal rates tend to produce lower short-run real and nominal exchange rates volatility, while variations in the two rates are almost perfectly correlated. The presence of price sluggishness is the most intuitive explanation for this result: the price level only adjusts slowly to changes in nominal exchange rates. In this case, variations in the nominal exchange rate may strongly affect the real exchange rate. Moreover, the whole transmission mechanism of aggregate supply and demand shocks may be different under alternative international monetary arrangements.

In general, empirical studies on the matter have estimated small structural (VAR) models of the economy, often in the spirit of the Fleming-Mundell-Dornbusch tradition, to analyse the importance of these considerations. In this vein, Kaminsky and Klein (1994) estimate a structural VAR to decompose the shocks to the dollar/pound real exchange rate during the Gold Standard. Their main finding is that shocks to fiscal deficits were associated with fluctuations of the exchange rate, while changes in government spending *per se* were not having a systematic effect on it.

Clarida and Gali (1994) perform a similar task, in turn, on the dollar real exchange rate with the other G3 countries over the post-Bretton Woods period. Their findings are that aggregate demand shocks in general had significant long-run effects on the real exchange rate, while monetary shocks tend to die out much quicker and account only for a small portion of its variability. Supply shocks, instead, play no role in real exchange rate dynamics. Eichenbaum and Evans' (1995) estimated model yields the same results.

MacDonald (1996) estimates a cointegrating VAR over the same period and countries. Impulse-response analysis carried out after imposing some long-run restrictions reveals that fiscal balance and productivity shocks do produce appreciation of the real exchange rate, but this is short-lived and quickly reversed[11].

Rogers (1999) is one of the two studies -the other being Muscatelli and Spinelli (1999)- that uses long historical data sets. It examines the relevance of various kinds of shocks on the dollar/pound real exchange rate between 1889 and 1992. A structural VAR analysis is conducted on a variety of system vectors including alternative measures of the fiscal stance, productivity levels, and money market indicators. The SVAR is then identified applying Blanchard and Quah's (1989) identification restrictions on the long-run response of variables to the shocks. The main result from this study is that, contrary to Clarida and Gali's (1994) findings, monetary shocks account for nearly half of the observed variability in the real exchange rate, though, again, these effects tend to die out relatively fast. In addition, fiscal and trade policies appear to have a substantial impact on the exchange rate, while productivity levels do not seem to be the a key determinant of the variance of real exchange rate changes. These results should be carefully borne in mind when studying impulse responses in our time-varying context[12].

Muscatelli and Spinelli (1999) estimate a semi-structural VAR model in the real exchange rate and deviations of domestic productivity levels, real ex ante interest rates and government spending as a fraction of GDP from their foreign counterparts. The study uses the same data set as ours (1870-1995), and it is then particularly relevant[13]. Identification of the structural shocks is achieved through a recursive ordering of the variables. Results confirm that in many cases PPP does not hold even over long time horizons. In producing

[11] Similar results are achieved by Clark and MacDonald (1998).

[12] However, some words of caution should be spent on Rogers' (1999) results. Impulse responses are presented within error bands representing the 84$^{th}$ and 16$^{th}$ percentiles of the empirical distribution of 1000 simulated impulse responses. This amounts to assuming a confidence level of about 70%.

[13] Compared to ours, their country sample is extended to include France.

such deviations fiscal and productivity shocks appear to have a substantial and persistent impact. In line with Rogers (1999), the authors find that a Balassa-Samuelson effect on real exchange rates can be tracked well before and into the post-Bretton Woods period.

Finally, Clarida and Prendergast (1999) re-examine the question for the dollar/G3 rates, using a structural VAR consisting of the OECD's estimates of the structural primary budget surplus relative to potential GDP, the output gap, the ratio of the actual primary budget surplus to actual GDP, and a trade-weighted index of the real exchange rate. Their findings are that the real exchange rate substantially appreciates in response to an expansionary fiscal shock, but again, this movement is eventually reversed after a few observations. Clarida and Prendergast's (1999) work is certainly a useful starting point for analyses devoted to the study of the dynamic effects of fiscal policy (see Fatas and Mihov, 1998, Blanchard and Perotti, 1999). However, the lack of a significant number of observations –their study employs annual data over 1975 to 1996- does not permit to take the above results as particularly robust[14].

### *2.4 Real Exchange Rates, Productivity, and Fiscal Policy Shocks: Where Do We Stand?*

We are now in the position to express some remarks on the main findings of the studies just examined.

a) There is some evidence pointing towards diverging productivity patterns as having some persistent effect on real exchange rates. This is in line with the BSH, though there is little available empirical evidence as to the relative importance of this productivity-trend effect in determining observed movements in the real exchange rates. Also, there is some hint about the possibility that such relationship, as it depends on changing relative productivity levels, may not be invariant over time.

b) Some additional evidence suggests that positive shocks to the fiscal balance might have some appreciating effect on the relative price of traded goods. Such a relationship, however, appears to be confined to short-run deviations from equilibrium, and in some studies its sign is not in line with theory's predictions.

c) The persistence profile of both kinds of shocks appears to be very controversial. This does not come as a surprise, however. All recent intertemporal macro models (Obstfeld and Rogoff, 1997) show that the persistence profile of such shocks is affected by changes in consumer preferences, demographic and technological factors, as well as by capital and labour mobility. The question is then an eminently empirical one, and it calls for some caution in employing structural models to disentangle the issue.

d) The variability in the results of many of the empirical analyses surveyed above also suggests that the data frequency, and the period over which such studies are conducted, are indeed crucial. The main reason might be that the transmission of monetary and non-monetary shocks is not invariant to the monetary regime in place, and that real exchange dynamics is better decomposed using low-frequency data.

## 3. VAR Models of Real Exchange Rate Fluctuations: Time-Invariant or Time-Varying?

In the light of the above considerations about the existing evidence, we decided to study the dynamic response of bilateral real exchange rates between the US, UK and Italy from a time-varying VAR perspective. Since the empirical methodology we apply deserves careful examination, our next steps are as follows. We first illustrate the way we cast the model illustrated in Section 2.1 into a standard VAR representation. Then, we turn to discussing how such form can be further developed into one that better takes into account the intrinsic time-varying nature of the relationships we are investigating.

Equation [14] below (in which the bar on $q$ emphasises the character of long-run, equilibrium value for the real exchange rate)

$$\bar{q}_{t+1} = -\alpha[\bar{a}^{T}_{t+1} - f_N \hat{a}^{N}_{t+1} + (f_N - 1)\hat{g}_{t+1} + \hat{p}_0] \qquad [14]$$

was derived in the context of a small structural model in which the basic features of the recent wave of intertemporal macro models were all taken into account. It

[14] This is likely to be the reason why the authors do not provide standard errors bands for their estimated impulse responses.

represents a relationship between the real exchange rate, relative productivity levels in the traded and nontraded sectors, and government spending as a proportion of GDP. This specification, however, illustrates a long-run, equilibrium relationship, and would naturally call for a cointegration approach[15] (MacDonald, 1996). However, Muscatelli and Spinelli (1999) have already attempted such route, and for the same exchange rate series as ours, found no robust evidence of cointegration. Furthermore, we believe that a cointegrating approach employed on a very long dataset as ours, would end up assuming a fixed relationship amongst the variables, which is partly contradicting our original aim. Finally, the points we raised at the end of last section suggest to avoid being too dogmatic about the relationship between structural models and econometric specifications for the real exchange rate.

Drawing upon the negative findings of Muscatelli and Spinelli (1999) about the joint trending properties of our time series, we then choose to re-formulate the above model. In particular, following MacDonald (1996) and, *inter alia*, Meese and Rogoff (1988), we now anchor a real exchange rate short-run adjustment mechanism to the long-run relationship in [14][16]. Consider the standard uncovered interest parity condition for a one-period bond:

$$E_t\left\{\Delta e_{t+1}\right\} = \hat{i}_t + \hat{u}_t^{m}, \qquad [15]$$

where $i$ is the nominal interest rate (again, circumflex denotes deviations from the foreign country's respective value) and $u^{m}$ is a relative money demand shock. Subtracting the expected inflation differential in $t + 1$ from both sides we obtain a measure of the real interest parity condition expressed in terms of the actual equilibrium exchange rate:

$$q_t = E_t\left\{q_{t+1}\right\} - \hat{r}_t, \qquad [16]$$

where $r_t = i_t - E_t\left\{\Delta p_{t+1}\right\}$ is the real *ex ante* interest rate. If we further assume that the first term on the RHS of [16] coincides with the equilibrium real exchange rate (in

[15] Please recall, however, the difficulties the literature finds in detecting such cointegrating relationship empirically. See for example Chinn (1997).

[16] Clark and MacDonald (1998) and MacDonald and Nagayasu (1999) find that Meese and Rogoff's (1988) inability to detect strong evidence of thwe long-run link between real exchange rates and real interest differential might be due to the estimation technique. MacDonald and Nagayasu (1999) find strong evidence in favour of such link by using panel cointegration techniques.

$t+1$) we have in [14], we derive an expression in which the actual real exchange rate is a function of productivity and fiscal policy fundamentals and the real interest differential:

$$q_{t+1} = \overline{q}_{t+1} - \widehat{r}_{t+1} = -\alpha[\widehat{a}^T_{t+1} - f_N\widehat{a}^N_{t+1} + (f_N - 1)\widehat{g}_{t+1} + \widehat{p}_0] - \widehat{r}_{t+1} \qquad [17]$$

The above equation is our final relationship. The presence of the real interest differential constitutes the main mechanism through which the real exchange rate adjusts towards its long-run equilibrium. It should be noticed, once again, that it is the absence of perfect intersectoral factor mobility that allows us to obtain persistent fluctuations of the real exchange rate.

Expression [17] can be trivially turned into the following unrestricted VAR($p$) specification[17]:

$$X_t = c + \sum_{j=1}^{p} A_j X_{t-j} + \varepsilon_t \qquad t = 1,2,\ldots,T;\ j = 1,2,\ldots,p, \qquad [18]$$

where $X_t$ is a $(n \times 1)$ vector of endogenous variables, $A_j$s are the $(n \times n)$ matrices of parameter coefficients, and $\varepsilon_t$ is a $(n \times 1)$ vector of disturbances for which, of course

$$E\{\varepsilon_t\} = 0$$
$$E\{\varepsilon_t \varepsilon_t^{'}\} = \Sigma$$
$$E\{\varepsilon_t \varepsilon_s^{'}\} = 0,\ \forall t \neq s$$

In our case, $X_t^{'} = \left(q, \widehat{\tau}, \widehat{D}, \widehat{r}\right)_t$ and $n = 4$. $q$ represents a CPI-based measure of the bilateral real exchange rate, $D$ is a measure of the fiscal imbalance as a fraction of GDP, and $\tau$ is the relative level of productivity in the traded sector[18]. As before, circumflexes indicate that variables are expressed as departures of domestic magnitudes from their respective foreign counterparts.

[17] We remind that Clarida and Gali (1994), Kaminsky and Klein (1994), and Muscatelli and Spinelli (1999) adopt semi-structural VARs. MacDonald (1996) estimates a (Johansen) cointegrating VAR and performs structural analysis on the orthogonalised shocks. The cointegrating relationship involves assuming that the real exchange rate depends on real interest rate differentials and the deviations of fiscal and productivity variables from foreign values. MacDonald's (1996) approach appears very close to ours, in that it assumes a long-run relationship similar to our [13] and adopts the same information set as in our subsequent time-varying VAR estimates.

Now, the existing empirical literature adopting VAR approaches (whether cointegrating or not) is confronted with the difficult task of devising some identifying restrictions in order to recover the fundamental disturbances from the VAR residuals, and to perform the necessary structural analysis. Following Lutkepohl (1991), Hamilton (1994), and Pesaran and Shin (1998), we re-cast the model in [18] in the following way[19]:

$$
\begin{aligned}
&X = AZ + U,\\
&X = \begin{pmatrix} X_{p+1} & X_{p+2} & \dots & X_T \end{pmatrix};\\
&A = \begin{pmatrix} cA_1 & \dots & A_p \end{pmatrix}; Z = \begin{pmatrix} Z_p & Z_{p+1} & \dots & Z_{T-1} \end{pmatrix}\\
&Z_t = \begin{pmatrix} 1 \\ X_t \\ X_{t-1} \\ \vdots \\ X_{t-p+1} \end{pmatrix}; U = \begin{pmatrix} \varepsilon_{p+1} & \varepsilon_{p-2} & \dots & \varepsilon_T \end{pmatrix}
\end{aligned}
\qquad [19]
$$

where now we have only $T^* = T - p$ observations available in each equation.

Assuming that the model is stationary is equivalent to saying that the VAR has the following finite MA representation

$$X_t = \sum_{j=0}^{\infty} B_j \varepsilon_{t-j}, \qquad [20]$$

where the $B_j$s are the $(n \times n)$ MA parameter matrices. Given the information set $\Omega$, if the residual variance-covariance matrix $\Sigma$ is diagonal, the impulse-response function will be defined as

$$IR_X\left(h, \delta, \Omega_{t-1}\right) = E\left\{X_{t+h} \middle| \varepsilon_t = \delta, \Omega_{t-1}\right\} - E\left\{X_{t+h} \middle| \Omega_{t-1}\right\}, \qquad [21]$$

namely, the difference between the expected value of $X_t$ at horizon $h$, given that a shock $\delta$ hits the system in time $t$, and the expected value of $X_t$ in the absence of shocks.

[18] Below we examine the key issues relative to the measurement of such variables in our historical context.
[19] Henceforth we assume that $p$ is known.

The MA parameters $\varphi$ can then be interpreted as responses of $X_{t+h}$ to a shock in $t$ on variable $j$:

$$\varphi_{j,h} = B_h e_j , \qquad [22]$$

where $e_j$ is a vector of zeroes with one as the $j$th element.

If, on the other hand, $\Sigma$ is not diagonal, contemporaneous interactions amongst the variables prevent any interpretation of the VAR residuals as fundamental disturbances, and the system is not identified.

The solutions put forward by the traditional VAR literature (Blanchard and Quah, 1989; Hamilton, 1994) run essentially along two main avenues. First, one might try to impose some structure on the VAR by orthogonalising the shocks according to a Choleski decomposition of the residual variance-covariance matrix: $PP' = \Sigma$, where $P$ is a lower triangular matrix. This way, the orthogonalised responses are recovered as

$$\varphi^{O}_{j,h} = B_h P e_j \qquad [23]$$

Of course, devising congruent identifying restrictions is a major task that presents the researcher with problematic dilemmas about the ordering of the variables in the system, which in turn determines the dynamic interactions amongst reduced-form residuals.

Alternatively, Blanchard and Quah (1989) proposed to impose restrictions on long-run MA parameter matrices only, leaving the short-run movements of variables untouched. As in the case of recursive-triangular representations of the system, however, this amounts to impose some structure on the variables in order to recover the fundamental shocks. This obviously entails making relatively stringent assumptions on the transmission of the shocks. There is no agreement on the ideal recursive ordering or identifying structure to impose to our set of variables. In particular, it would be peculiarly hard to imagine a clear-cut way of configuring the apparent correlation existing between productivity and fiscal variables. Moreover, in our particular case, a fixed recursive ordering or a method à la Blanchard-Quah, would also entail that the transmission of the shocks is imposed to be invariant to fundamental changes occurring in the economy. Since we employ a relatively long-spanning data set, the latter idea sounds particularly odd. The approach we use is instead inspired to avoiding such restrictions. This is why we chose to follow, for the structural analysis following the time-varying estimation of our VAR systems, Koop, Pesaran and Potter's (1996) and Pesaran and Shin's (1998) Generalised Impulse Response methodology. They propose to employ an approach in which the empirical distribution of the response to a large number of different shocks to the system is examined. In our case, this approach will allow to overcome considerable difficulties we would meet in envisaging the interaction between the variables, and in attempting to identify the transmission of the shocks. This is particularly helpful, also because the contemporaneous correlations between productivity and fiscal stance measures are likely to be substantial.

Let us consider the impulse response of $X_{t+h}$ to a shock on variable $j$, assuming that there are contemporaneous interactions amongst the variables:

$$\begin{aligned} IR_X\left(h,\delta_j,\Omega_{t-1}\right) &= E\left\{X_{t+h}\left|\varepsilon_{t,j}=\delta_j,\Omega_{t-1}\right.\right\}-E\left\{X_{t+h}\left|\Omega_{t-1}\right.\right\}= \\ &= B_h E\left\{\varepsilon_t\left|\varepsilon_{t,j}=\delta_j\right.\right\} \end{aligned} \qquad [24]$$

If the distribution of $\varepsilon_t$ is multivariate normal, the conditional expectation of $\varepsilon_{t,j}$ given the shock in equation $j$ is

$$E\left\{\varepsilon_{t,i}\left|\varepsilon_{t,j}=\delta_j\right.\right\}=\frac{\sigma_{ij}}{\sigma_{jj}}\delta_j, \qquad [25]$$

where the $\sigma$s are elements from $\Sigma$. Accordingly, the *generalised* impulse response is defined as

$$IR_X\left(h,\delta_j,\Omega_{t-1}\right)=B_h\begin{pmatrix}\sigma_{1j}\\ \vdots\\ \sigma_{jj}\\ \vdots\\ \sigma_{nj}\end{pmatrix}\frac{\delta_j}{\sigma_{jj}}=B_h\Sigma e_j\frac{\delta_j}{\sigma_{jj}}=\frac{B_h\Sigma e_j}{\sqrt{\sigma_{jj}}}\frac{\delta_j}{\sqrt{\sigma_{jj}}}=$$
$$=\frac{B_h\Sigma e_j}{\sqrt{\sigma_{jj}}}=\varphi_{j,h}^{G} \qquad [26]$$

In other words, the generalised impulse response function is based on an "average" shock hitting the system. It is equivalent to the function generated through a conventional Choleski decomposition when the residual variance-covariance matrix is diagonal.

In the same vein, defining the *MSE* matrix of a $h$-step forecast of $X_t$ by

$$MSE\left(X_t(h)\right)=\sum_{j=0}^{h-1}B_j\Sigma B_j', \qquad [27]$$

the error of a $h$-step forecast of a variable $X_{k,t}$ can be decomposed into the contributions of innovations to the variables of the system. In the case of orthogonalised and generalised responses, we have, respectively, the following forecast –error variance decomposition functions:

$$\theta_{k,i}^{O}(h)=\frac{\sum_{j=0}^{h-1}\left(e_k'B_jPe_i\right)^2}{MSE\left(X_{k,t}(h)\right)}$$
$$\theta_{k,i}^{G}(h)=\frac{\frac{1}{\sigma_{kk}}\sum_{j=0}^{h-1}\left(e_k'B_j\Sigma e_i\right)^2}{MSE\left(X_{k,t}(h)\right)} \qquad [28]$$

An important caveat of this approach is that the interpretation of impulse responses and forecast error variance decomposition is different from the case of othogonalised disturbances. If the residual variance-covariance matrix is not diagonal, innovation accounting is such that the sum of all contributions to explaining, for example, the total forecast error variance, does not add up to one. This means that comparisons over which variables contribute most to such variance cannot be conducted in percentage levels, but in relative terms.

Amongst the contributions we recurrently refer to[20], Rogers (1999) employs Blanchard-Quah's structural-VAR framework. MacDonald (1996) estimates a cointegrating VAR on which long-run restrictions are imposed to investigate the impulse responses of the real exchange rate to various orthogonalised shocks to VAR's elements. Muscatelli and Spinelli (1999) use for the real interest rate and productivity differentials, and for the real exchange rates, the same data set as ours. No cointegrating relationship is in turn found. The authors then identify the fundamental disturbances to the system through some restrictions imposed on the short-run parameter matrix.

Against this background, it is clear that the VAR approach overall represents quite a natural instrument to examine the relative importance of the fluctuations we are investigating. However, we have sought to circumvent the problematic identification tasks sketched above, not only by conducting structural analysis according to a generalised impulse response framework, but also by taking a radically different route as far as the VAR estimation is concerned. Following a recently developed strand of literature, we adopt a Bayesian perspective to the specification of the VAR system (Canova, 1995; Hamilton, 1994). We try to "enhance" the signal-to-noise content of the historical data set we use by adopting "*....a symmetric atheoretical prior on all the variables to trade off overparameterization with oversimplification*"[21]. We first cast our baseline model into a convenient state-space representation, and then apply the Kalman filter to recursively compute an optimal estimate of the particular unobserved state vector used to explain the system's dynamics. Our formulation then allows the VAR coefficients to be time varying, and, as we will see in the next Section, proposes a peculiarly unrestricted dynamics for the state vector. This avoids resorting to strict identifying restrictions and estimation procedures that might bias the economic hypotheses we are testing for[22].

Our model is close to a number of time-varying coefficient models, developed during the eighties, which devoted particular attention to the specification of optimal probability distributions for the coefficients (Nicholls and Quinn, 1982; Quinn, 1986). Also, our approach is close, at least in spirit, to the studies stemming from the classic

[20] See Section 2.3
[21] Canova (1995), p. 85.
[22] Bec, Ben Salem and MacDonald (1999) estimate a Threshold Autoregressive model (TAR) of seven real exchange rates in which the thresholds are defined in terms of the real interest rate differential. Substantial nonlinearities are then

Bayesian analysis of Doan, Litterman and Sims (1984). Our investigation differs from those analyses in that we do not rely on any prior belief about the autoregressive process of each time series. This will be particularly evident in the next section, where we clarify that the "atheoretical" prior we adopt is particularly "loose", and characterised by the smallest possible number of parameters. Furthermore, in contrast with standard structural-VAR approaches, the time-varying specification adopted here does not require taking into consideration the eventual trending behaviour of some or all the variables included in the system. The analysis hinges on the likelihood principle, which makes it unaffected by the presence of unit roots (Canova, 1995).

The search for the optimal definition of the state vector is performed by applying the Kalman filter recursively over the sample. This yields (Harvey, 1989), under the assumption of normality of the residuals, the parameters (prediction error and its variance) of the optimal prediction error decomposition of the likelihood function. In turn, the selection of the optimal state vector is carried out numerically, through evaluation of the likelihood function. Structural analysis is finally performed over each observation of the state-space VAR model.

The dynamic response of the real exchange rate to shocks in each of the remaining variables of the system is finally examined, using the Generalised Impulse Response Approach of Koop *et al.* (1996) and Pesaran and Shin (1998). In principle, one possible alternative to this route would have involved assuming a particular recursive ordering of the variables in the VAR, and perform structural analysis through Choleski decomposition of the residual variance-covariance matrix. As already discussed, this approach was followed by Muscatelli and Spinelli (1999). We believe that, given the long span of data, and the characteristics of the variables included in the model, an appropriate and unquestionable identification scheme for the various shocks within the system would be hard to find. As we have seen, despite the recent progress in the development of open economy intertemporal models, there is still wide disagreement as to the ways in which, for example, monetary and real shocks determine real exchange ate fluctuations. More specifically, we suspect that the degree of correlation between fiscal and productivity variables goes well beyond any effort to identify the transmission of shocks between them. Finding a recursive structure that fully accounts for such correlations would certainly represent a daunting task. Furthermore, a fixed recursive scheme assumed for the

detected, indicating that real fundamentals, like the ones considered in our work, might be very relevant in explaining transmission of fundamental shocks to the variables would amount to treating the 130 years of data as generated under one single monetary regime. This would be in open contrast with the aim of unveiling whether and how the transmission of shocks has changed over time.

The next Section illustrates in more details our approach.

## 4. A Time-Varying VAR Perspective on Real Exchange Rates

A transition and a measurement equation overall define a state-space model (see Harvey, 1989; Hamilton, 1994; Kim and Nelson, 1999). The **measurement equation** describes the dynamic relationship the model postulates between a $n \times T$ vector of observable variables $y_t$, and the so-called **state vector** $\beta_t$. The latter is usually assumed as generated by a first-order Markov process[23]:

$$\beta_t = T_t \beta_{t-1} + c_t + R_t \eta_t \qquad t = 1,...,T \qquad [29]$$

where $\beta_t$ is a $m \times 1$ vector, $T_t$ is a $m \times m$ matrix, $c_t$ is a $m \times 1$ vector, and $R_t$ is a $m \times G$ matrix. The measurement equation is in turn:

$$y_t = Z_t \beta_t + d_t + \varepsilon_t \qquad t = 1,...,T \qquad [30]$$

where $Z_t$ is a $n \times n$ matrix that links the observed variable $y_t$ and the state variable $\beta_t$, and $d_t$ is a $n \times 1$ vector. The disturbances $\eta_t$ and $\varepsilon_t$ are white-noise, uncorrelated processes of dimension $G \times 1$ and $n \times 1$, respectively, with time-varying covariance matrices:

$$\begin{pmatrix} \eta_t \\ \varepsilon_t \end{pmatrix} \sim MVN \begin{pmatrix} 0 & Q_t \\ 0 & H_t \end{pmatrix}$$

Once a model is formulated in state-space form, the Kalman filter is initially applied to obtain estimates of the unobserved state vector conditional on some starting

long-run real exchange rate movements. Bayoumi and MacDonald's (1998) conclusions point to the same direction.

value for the state vector and its variance. In correspondence of each observation, when system residuals are normally distributed, an evaluation of the log-likelihood function based on the prediction error decomposition yields an intuitive way of updating those starting values. Iterating the filtering procedure from $t = 1$, and evaluating the log-likelihood function from $t + 1$ onwards, minimises the effect of the starting values, and yields estimates of the state vector based on information available up to time $t$[24].

Now, if a new observation on the data is available at the end of each period, the basic Kalman filter consists of two steps: **prediction and updating**[25]. In the former, an expectation of the state vector $\beta_t$ conditional on information up to the previous period is computed. Subsequently, a forecast of $y_t$ is calculated, along with the prediction error. Since this contains information about the state vector that goes beyond that contained in the estimated value for the previous period, a more accurate inference about it can now be made, based on information up to the present period. Using such information, the estimate of the state vector is updated.

Analytically:

| **Prediction** | **Updating** | |
|---|---|---|
| $\beta_{t\|t-1} = T_t\beta_{t-1} + c_t$ | $\beta_t = \beta_{t\|t-1} + K_t v_t$ | |
| $P_{t\|t-1} = T_t P_{t-1} T_t' + R_t Q_t R_t'$ | $P_t = (I_m - K_t Z_t) P_{t\|t-1}$ | [31] |
| $y_{t\|t-1} = Z_t \beta_{t\|t-1} + d_t$ | where | |
| $v_t = y_t - y_{t\|t-1}$ | $K_t = P_{t\|t-1} Z_t' F_t^{-1}$ | |
| $F_t = Z_t P_{t\|t-1} Z_t' + H_t$ | | |

[23] If $F$ is a diagonal matrix with values less than one as its diagonal elements, the state vector follows a stationary autoregressive process.
[24] Along with this *basic filter*, the *smoothing* procedure yields estimates of the state vector based on all the available information in the sample up to $T$.
[25] For more details about the derivation of these two steps, see Harvey (1989) or Hamilton (1994).

In the above, $P$ is the covariance matrix of the state vector, and $K$ represents the Kalman gain, which determines the weight assigned to new information contained in the prediction error. We initialise the recursive procedure by estimating an unrestricted VAR in $X_t = \left(q, \tilde{\tau}, \dot{D}, \tilde{r}\right)_t$. Standard lag-length pre-tests are applied, yielding $p = 3$ and $p = 2$, depending on the pair of countries under examination, as the optimal order. Next, we build the initial state-vector using the VAR estimated values of the residual variance-covariance matrix for each equation and all coefficients' standard errors. The model is then cast into the state-space formulation we shortly illustrate below, and the (basic) Kalman filter procedure is passed recursively through the sample[26]. This way, we produce an optimal estimate of a first state vector, providing values, *inter alia*, for the residual variance of each equation. Simultaneously, we apply a second Kalman filtering procedure -this time with an embedded smoothing algorithm- on the VAR's coefficients. A vector of estimated coefficients is then generated for each observation. This way, we allow the contribution of single variables to the explanation of the total variance of each equation to vary in correspondence of each observation.

In time-varying formulations of the state-space model, the Z matrix is represented by a matrix of exogenous and/or predetermined variables. This is exactly what happens in our model (recall that our baseline specification involves $n = 4$ and $p = 2$ or 3), in which lags of the endogenous variables are included in the Z matrix:

$$Z_t = \begin{bmatrix} y_{1,t} & 0 & 0 & y_{2,t} & 0 & 0 & & y_{n,t} & 0 & 0 \\ 0 & y_{1,t} & 0 & 0 & y_{2,t} & 0 & \cdots & 0 & y_{n,t} & 0 \\ 0 & 0 & y_{1,t} & 0 & 0 & y_{2,t} & & 0 & 0 & y_{n,t} \\ & \vdots & & & \vdots & & \ddots & & \vdots & \\ y_{1,t-p+1} & 0 & 0 & y_{2,t-p+1} & 0 & 0 & & y_{n,t-p+1} & 0 & 0 \\ 0 & y_{1,t-p+1} & 0 & 0 & y_{2,t-p+1} & 0 & \cdots & 0 & y_{n,t-p+1} & 0 \\ 0 & 0 & y_{1,t-p+1} & 0 & 0 & y_{2,t-p+1} & & 0 & 0 & y_{n,t-p+1} \end{bmatrix} \quad [32]$$

[26] Maximum Likelihood Estimation (BFGS algorithm) is used. All estimates were produced using a GAUSS code purposely elaborated with Ulrich Woitek, whom I am particularly grateful to. The code is available from the author upon request. The Kalman filter and ML procedures employed TSM, an advanced time-series package for GAUSS, by Thierry Roncalli.

Furthermore, we assume that all the elements of the $T_t$ matrix, and of the $c_t$ vector (refer to equations [29]-[30]), are zeroes. This amounts to hypothesise that the state vector of the model behaves according to a white-noise stationary process. The rest of our state-space model is as follows

$$\begin{aligned} R_t &= I_{p^* n^2} \\ d_t &= \begin{bmatrix} \beta_{1t} & \beta_{2t} & \beta_{3t} & \beta_{4t} \end{bmatrix}' \end{aligned} \qquad [33]$$

whereas the $Q$ and $H$ matrices are made up of appropriately chosen elements of the state vector. To sum up, the state-space representation we adopt is, in terms of equations [29]-[30], the following:

$$\begin{aligned} y_t &= Z_t \beta_t + d_t + \varepsilon_t \qquad t = 1,...,T^* \\ \beta_t &= R_t \eta_t \qquad t = 1,...,T^* \end{aligned} \qquad [34]$$

One appealing feature of the state space representation we adopt is that, by allowing the behaviour of the state vector to be unaffected by its past values, we avoid a major risk of using historical data. That is, we prevent the volatility of observations to be exceedingly reflected in the estimation of the state vector. Such a case would prevent us from capturing the many of the shocks of our system, as the signal-to-noise ratio perceived in the case of, say, a stationary autoregressive process for the betas in [34], would probably be lower.

## 5. Investigating Real Exchange Rate Fluctuations

Before evaluating the dynamics of the three real exchange rates we study, we perform a series of standard Augmented Dickey-Fuller tests for nonstationarity on all our series. For each country, we pre-test for the level of integration of each of the elements in the system vector $X_t = \left(q, \tilde{\tau}, \tilde{D}, \tilde{r}\right)_t'$. Of course, Augmented Dickey-Fuller tests are not the optimal way of assessing the stationarity properties of our series. Already Froot and Rogoff (1995) forcefully argued in favour of cointegration tests rather than standard tests for the mean reversion of real exchange rates, on the grounds that the former provide less stringent assumptions concerning the relative price of nontraded goods. Accordingly, Muscatelli and Spinelli (1999) perform, *inter alia*, Johansen cointegration tests on the same (CPI-based) real exchange rate series we study, finding very mixed results. For the US dollar/UK sterling real exchange rate, one cointegrating vector is identified, but the unit restriction on the coefficients on the price levels is rejected. For the Italian lira/US dollar rate there is more evidence of cointegration, while in the case of Italian lira/UK sterling rate, no significant eigenvectors are found.

More generally, a number of recently devised alternative tests, with trend-stationarity as a null (Kwiatowski, Phillips, Schmidt and Shin (1992), for example), or which account for structural breaks in the trend, are likely to be better suited to investigate such properties[27]. It is also clear that a system perspective, where the effects of temporary and permanent shocks of endogenous variables are more closely pinned down, is a much better framework for conducting a study of the trending properties of real exchange rates. However, a cointegrated-VAR over a long historical data set would prove inappropriate, as the long-run relationships of both productivity and fiscal policy factors on the exchange rate may well change over time. This is one of the eye-catching features displayed by the three series, as plotted, for each country, in **Figures 5** to **7** (see below a description of the way we measure each differential). Since the subsequent time varying VAR analysis is unaffected by the cointegration or nonstationarity properties of the series under investigation, in our context this issue does not appear to be of primary relevance.

The real exchange rate series we use are CPI-based (Consumer Price Index) measures of the bilateral real exchange rate. Given the long time span covered by the data, we expect to find, in the ADF tests, some evidence supporting stationarity. The bilateral exchange rate series are shown in **Figure 1**, where we have scaled means and ranges to allow better visual impression of the movements in the bilateral rates. The real exchange rate series have been constructed in a way that a decrease in their value constitutes a real appreciation.

**Table 1** below displays results from simple ADF tests on the three bilateral exchange rates, using both CPI and WPI (Wholesale Price Index) definitions.

[27] See Maddala and Kim (1998) for an excellent survey of such methods. However, Engel (1996) shows that tests, like those in Kwiatowski et al. (1992), suffer from severe shortcomings when applied to real exchange rate series. In addition, such tests appear biased towards finding stationarity, exactly as in the case of traditional ADF tests.

| Test/Bil. Ex. Rate | CPI Definition | | WPI Definition | |
|---|---|---|---|---|
| | $t_\mu$ | $t_\tau$ | $t_\mu$ | $t_\tau$ |
| ***UK/US*** | -3.1307* | -3.7242* | -3.3423* | -3.3998 |
| ***IT/US*** | -4.5157** | -4.7174** | -3.7311** | -3.8711* |
| ***UK/IT*** | -2.7417 | -3.2759 | -3.0487* | -2.9505 |

**Table 1** – Real Exchange Rates, Augmented Dickey-Fuller tests. Sample: UK/US, 1861-1995; IT/US, 1861-1996; UK/IT, 1861-1995. Real exchange rates are defined using the Consumer Price Index (CPI) and the Wholesale Price Index (WPI). The numbers in the columns labelled $t_\mu$ and $t_\tau$ refer to ADF $t$-ratios from regressions (two lags were employed) in which a constant only and a constant plus a trend, respectively, were included. "*" and "**" indicate that the null of non-stationarity is rejected at the 5% and 1% significance level, respectively. Critical values are from Banerjee, Dolado and Mestre (1992)

The test statistics are computed from autoregressions with a constant only, and with a constant and a time trend. In most cases, the null of nonstationarity is rejected, with the notable exception of the CPI-defined UK/Italy exchange rate. The fact that with WPI-based measures of the real exchange rate the null of nonstationarity tends to be rejected more easily than with equivalent CPI-based series, is well known in empirical work[28], as well as the strong sample-dependence of the results[29].

Next, we test for the level of integration of the series we use for the productivity, budget position, and real interest rate differentials. All differentials are defined as home values minus respective foreign country's values.

Productivity levels are proxied by real GDP per capita. Obviously, there are important shortcomings associated with such a measure, while output per worker could have represented a better indicator. However, given the long time span and the lack of reliable data on employment and capital stocks, our choice is rather limited.

Our measure of the fiscal imbalance is the ratio of government spending minus tax revenues on real GDP. Muscatelli and Spinelli (1999) adopt government spending on GDP as their measure of the fiscal stance. This, however, does not allow capturing the effect of tax-driven changes in the budgetary position. Of course, it is not clear how tax-driven fiscal shocks –as opposed to spending shocks- should be interpreted in terms of the Balassa-Samuelson hypothesis and, more generally, of the intertemporal approach to the

[28] See, for example, Muscatelli and Spinelli (1999). Bayoumi and MacDonald (1999), using a shorter sample, detect the same behaviour in their series. This is probably due to the smaller weight nontraded goods have in WPI series.
[29] Rogers (1999), using a shorter sample, finds evidence of nonstationarity in the US/UK real exchange rate.

current account and the exchange rate. However, we do believe it is necessary to take into account a more appropriate measure of a country's fiscal position in our estimates. Unfortunately, data constraints do not allow us to take into account the effect of interest payments and transfers on gross government spending.

Finally, the real interest rate is computed as the nominal rate on long-term bonds minus the inflation rate in the previous period. **Figures 2** to **4** compare such series for the three countries.

**Table 2** below lists results from ADF test on the three series for each country. There is strong evidence of nonstationarity in the productivity differentials, whereas equally strong evidence of stationarity is found in the real interest differentials. The former result is in line with what one expects to find in a group of countries that have experienced relatively comparable productivity levels only recently by historical terms. On the other hand, the stationarity of the real interest differential is an intuitive and familiar finding (Bec, Salem, and MacDonald, 1999, for example). As regards the deviations in the fiscal indicators, results are less clear-cut. This can however be explained if one takes into account the impressive relevance of the war periods on the overall Italian-US differential series (and to a similar extent on the UK-Italian one), as displayed in **Figure 3**.

| *Test/ Country Pair* | ***Productivity differential*** | | ***Budget deficit/GDP ratio differential*** | | ***Real interest rate differential*** | |
|---|---|---|---|---|---|---|
| | $t_\mu$ | $t_\tau$ | $t_\mu$ | $t_\tau$ | $t_\mu$ | $t_\tau$ |
| ***UK/US*** | -2.5205 | -2.4313 | -3.8077** | -4.3647** | -6.5762** | -6.8507** |
| ***IT/US*** | -2.4404 | -2.2514 | -2.8132 | -2.8011 | -4.2307** | -4.2103** |
| ***UK/IT*** | -1.0772 | -2.4613 | -3.3052* | -3.3391 | -4.1271** | -4.1082** |

**Table 2** – Productivity, budget deficit/GDP ratio and real interest rate differentials, Augmented Dickey-Fuller tests. Sample: UK/US, 1871-1995; IT/US, 1867-1996; UK/IT, 1871-1995. Productivity levels are proxied by (the log of) GDP per capita levels. Budget deficit/GDP ratios are defined are government spending minus taxes over real GDP. The real interest rate is computed as the nominal rate on long-term bonds minus the inflation rate in the previous period. Differentials are defined as home values minus respective foreign country's values. The numbers in the columns labelled $t_\mu$ and $t_\tau$ refer to ADF $t$-ratios from regressions (two lags were employed) in which a constant only and a constant plus a trend, respectively, were included. "*" and "**" indicate that the null of non-stationarity is rejected at the 5% and 1% significance level, respectively. Critical values are from Banerjee, Dolado and Mestre (1992)

We now turn to our time varying estimates. For each country pair, we initially estimate the baseline VAR model consisting of the vector $X_t = \left(q, \hat{\tau}, \tilde{D}, \hat{r}\right)_t$ over the longest sample we have[30]. We remind that our approach lets the time-varying VAR coefficients contribute to explain each equation's variance, in turn optimally estimated using the Kalman filter, in a different way at each observation. This way, we allow the contribution of single VAR parameters towards explaining the total autoregression variance to change in correspondence of each year. Subsequently, we use generalised impulse responses to make some inference about the source of fluctuations in the real exchange rate. Impulse responses are bound between 95% confidence bands obtained through 500 Monte Carlo simulations. Analytical standard errors were in principle available (Pesaran and Shin, 1998), but are not sufficiently tested. On the other hand, bootstrapping (Davidson and MacKinnon, 1993) is a perfectly legitimate alternative to simulation, and in fact, it tends to yield roughly the same confidence bands. Despite finding a substantial number of insignificant responses at various horizons, we feel that our calculation of confidence bands took place along much more conservative grounds than in many related studies[31]. Though this will not modify the lines along which inference is conducted, we believe that the evaluation of our empirical exercise should give this some consideration.

**Figures 8** to **16** collect generalised impulse responses of the three bilateral exchange rates (RERATE) to shocks in the following differentials: productivity levels (PRODDEV), budget deficit/GDP ratios (BDDEV) and real interest rates (RDEV). The first column of each figure displays the effects on RERATE of its own shocks; the second, third and fourth columns, to shocks in each of the above variables, respectively. Each row shows impulse responses for selected years during the classical Gold Standard, the second post-war period, Bretton Woods and the free-floating experiences.

Starting with the UK sterling/US dollar exchange rate, the two panels in **Figure 8** show that shocks to the relative fiscal position do not have significant impacts on the real exchange rate. Shocks to the real interest rate differentials have some very limited depreciating effects, but these are negligible and very short-lived. On the contrary, under the classical Gold Standard, shocks to the productivity differential appear to have a

[30] UK sterling/US dollar, 1871-1995; Italian lira/US dollar, 1867-1996; UK sterling Italian lira, 1871-1995. Standard lag-length selection procedures were used to compute the optimal order of the VAR, found to be $p = 2$, or $p = 3$, depending on the country and the sample.
[31] Clarida and Prendergast (1999), for example.

significant, but not very persistent, depreciating effect. As a matter of fact, such effect becomes insignificant soon after the first year, when it does not reverse into a real appreciation (see third row in the Figure). On the other hand, the picture becomes less clear-cut in the second post-war period, with exchange rates first fixed and then allowed to float under the Bretton Woods and the post-1971 regimes (**Figure 8, Panel b**). The sterling real rate response to shocks in the productivity differential seems to become overall more persistent, and episodes of real appreciation alternate with ones of real depreciation. A cursory look at the forecast error variance decomposition reveals that, for the horizon at which all shocks seem to somewhat have some effect on the real exchange rate (h = 1), the contribution of innovations to the productivity differential is by far the largest[32]:

FEVD UK sterling/US dollar (typical year)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.94553 | 0.001311 | 0.026174 |

When we turn to analysing the Italian lira/US dollar case (**Figure 9**), we find that positive innovations to the productivity differential do have an appreciating effect on the real exchange rate, but this is apparently confined to the first year or so following the shock, with no particular differences across monetary regimes. When it comes to assessing the impact of shocks to the relative fiscal position, we observe that these tend to have a short-lived depreciating effect during the Gold Standard years, with the remarkable exception of the inter-war period[33]. In very recent years, however, in front of huge primary deficits accumulated by Italy since the late eighties, this pattern appears to be completely reversed: the initial depreciation is followed by a persistent real appreciation. The same changing pattern, but with an opposite sign, is displayed by the dynamic effects of a shocks to the real interest differential: these trigger a real appreciation during the Gold Standard years, and a persistent depreciation in more recent times[34]. This suggests to further investigate the links between shocks to the fiscal stance, net of interest payments, and the behaviour of interest rates. The relative contributions of the various shocks to

[32] Pesaran and Shin (1998) explain in detail why forecast error variance decomposition with generalised impulse responses does not yield percentage contributions. This is why interpreting forecast error variance decomposition functions in our context is much harder.

[33] After the suspension of convertibility Italy lived after 1913, the country briefly returned to the Gold Standard between 1927-1930.

[34] In recent years -between 1988 and 1995- such differential returned to be positive.

explaining real exchange rate's error variance, looks stable across the regimes covered by these estimates:

FEVD Italian lira/US dollar (1906)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.621211 | 0.086472 | 0.280335 |
| H = 2 | 0.618922 | 0.089069 | 0.283007 |
| H = 3 | 0.603851 | 0.111271 | 0.293212 |
| H = 4 | 0.596338 | 0.122271 | 0.298713 |
| H = 5 | 0.593059 | 0.127075 | 0.301194 |

FEVD Italian lira/US dollar (1927)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.621211 | 0.086472 | 0.280335 |
| H = 2 | 0.614344 | 0.093491 | 0.28827 |
| H = 3 | 0.502572 | 0.258735 | 0.36552 |
| H = 4 | 0.448099 | 0.338283 | 0.398392 |
| H = 5 | 0.388491 | 0.425896 | 0.449482 |

FEVD Italian lira/US dollar (1955)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.621211 | 0.086472 | 0.280335 |
| H = 2 | 0.621035 | 0.086681 | 0.280542 |
| H = 3 | 0.567779 | 0.165231 | 0.319632 |
| H = 4 | 0.561744 | 0.174058 | 0.32398 |
| H = 5 | 0.559807 | 0.176898 | 0.325389 |

FEVD Italian lira/US dollar (1990)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.621211 | 0.086472 | 0.280335 |
| H = 2 | 0.621075 | 0.086647 | 0.280496 |
| H = 3 | 0.491021 | 0.278475 | 0.377345 |
| H = 4 | 0.490974 | 0.278544 | 0.377382 |
| H = 5 | 0.49097 | 0.278549 | 0.377385 |

The figures above suggest that innovations to the productivity differential account for most of the variation in the forecast variance of the exchange rate at short horizons. At longer horizons, shocks to the remaining differentials become more relevant, with the real interest differential especially important in recent years, probably as a result of the increased international integration of Italian financial markets.

Next, we examine the UK sterling/Italian lira exchange rate. The picture emerging from **Figure 10** is not particularly clear-cut. The wide swings in the relative fiscal positions between the two countries introduce a stubborn amount of noise into the estimates, and make it peculiarly hard to assess the persistence of some effects. Shocks to relative productivity levels and to the real interest differential always have an appreciating effect on the exchange rate, as expected. This pattern, at least at the present stage, does not seem to be affected by the monetary regime in place. The relative importance of the various shocks to the decomposition of forecast error variance does not change much either across time:

FEVD UK sterling/Italian lira (typical year)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.891545 | 0.003553 | 0.347798 |
| H = 2 | 0.891383 | 0.003672 | 0.347733 |
| H = 3 | 0.887798 | 0.008344 | 0.347757 |
| H = 4 | 0.887095 | 0.009246 | 0.347742 |
| H = 5 | 0.886955 | 0.009427 | 0.34774 |

This also points to shocks in the productivity differentials and, to a lesser extent, in the real interest differential, as the main sources of innovation.

The main result from our estimates, so far, is that the evidence in favour of a persistent, productivity effect on the real exchange rate, if any, appears very thin. Instead, the link between fiscal shocks and real appreciation appears much more robust.

Only partly satisfied by the precision of our findings, we decided to conduct a further battery of estimates. That is, we split the initial sample into two sub-samples. The first covers the classical Gold Standard, running from around 1870 up to the First World War, whereas the second spans the Bretton Woods years plus the post-1971 floating experience. This way, we seek to introduce further room for variation in our estimates, allowing the VAR variance-covariance matrix to change across the historical data set we employ. The explicit cost of such strategy will involve giving up any attempt of using inter-war data for our estimates. However, a bird's eye view to the series in **Figures 1-7** suggests that such cost is not too high, compared to the level of noise brought by those observations.

**Figure 11, Panel a** illustrates generalised impulse responses for the sterling/dollar exchange rate from the model estimated over 1871-1913. The response of the real exchange rate in a typical year of the Gold Standard (1899) substantially confirms the irrelevance of the shocks to the relative fiscal position we had already found. In addition, shocks to the real interest differential appear irrelevant at all horizons. This time, however, the latter seems to trigger a real appreciation, in line with the predictions of our theoretical model, rather than depreciation. Shocks to the productivity differential, again, have a significant and very strong depreciating impact on the exchange rate, but not beyond the first year or so.

Estimation of the baseline model over the post-war sample yields the typical impulse responses illustrated in **Figure 11, Panel b**. The picture shows that during those years a shock to Britain's relative productivity levels did produce a short-lived depreciation of the exchange rate, of a much smaller size than during the Gold Standard. Innovations to the relative budget position instead trigger a real appreciation, this short-lived too. No persistent effects on the real exchange rate are detected.

When sub-sample estimation is conducted on the lira/dollar rate, the generalised impulse responses we obtain are displayed in **Figure 12**. **Panel a** shows results for the Gold Standard sample; in common to the full sample estimates they indicate that a shock to the relative budget deficit indicator leads the exchange rate to depreciate in real terms. On the contrary, a shock to productivity differential depreciates the rate, and by a remarkable extent[35]. Interestingly, such peculiar result is reversed in the post-war estimates (**Panel b**). The forecast error variance decomposition for a typical year of the sample confirms the increasing role of fiscal and monetary shocks over the latter part of the sample. Though not persistent, shocks originated in the demand side of the economy again appear to be of increasing relevance:

FEVD Italian lira/US dollar, 1867-1913 (typical year)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.601211 | 0.077472 | 0.240335 |

FEVD Italian lira/US dollar, 1952-1996 (typical year)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.984994 | 0.291946 | 0.681189 |

Finally, **Figure 13** illustrates our findings when system estimates are computed over the two sub-samples for the sterling/lira model. During the Gold Standard (**Panel a**), productivity and fiscal policy shocks seem to impact on the real exchange rate according to the BSH, whereas the response to a real interest differential shock is somehow perverse. The picture is not much different for the second post-war period (**Panel b**), where

[35] We have already detected a similar pattern with the UK sterling/US dollar rate.

however shocks to fiscal policy appear to have more enduring -though smaller- appreciating effects on the real exchange rate. **Panel c** also shows that in the free-floating period following the breakdown of Bretton Woods, in some instances the initial real appreciation is subsequently reversed into a real depreciation, and that the overall response tends to be more persistent. Forecast error variance decomposition for selected years confirms this finding, with post-1971 innovations in relative fiscal policy stance accounting for a noticeable larger amount of variation in the exchange rate:

FEVD UK sterling/Italian lira, 1949-1995 (1968)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.830231 | 0.125634 | 0.990171 |
| H = 2 | 0.830231 | 0.125634 | 0.990171 |
| H = 3 | 0.78362 | 0.179742 | 0.934102 |
| H = 4 | 0.757065 | 0.210565 | 0.902159 |
| H = 5 | 0.757064 | 0.210566 | 0.902158 |

FEVD UK sterling/Italian lira, 1949-1995 (1982)

| horizon | PRODDEV → RERATE | BDDEV → RERATE | RDEV → RERATE |
|---|---|---|---|
| H = 1 | 0.830231 | 0.125634 | 0.990171 |
| H = 2 | 0.830231 | 0.125634 | 0.990171 |
| H = 3 | 0.759194 | 0.208084 | 0.904721 |
| H = 4 | 0.72756 | 0.244802 | 0.866668 |
| H = 5 | 0.727559 | 0.244803 | 0.866667 |

Overall, the results we obtained using the sub-sample approach are interesting, because they tend to confirm that the response of the real exchange rate to shocks in productivity levels and the relative fiscal position is essentially sample-dependent. Again, such feature could have not been captured using a time-variant, fixed sample framework.

Moreover, it seems that such changing pattern is not strictly due to the exchange rate regime in place, as substantial differences emerge, within the post-war sample, in the responses obtained during and after the breakdown of Bretton Woods. Overall, we found relatively little evidence in favour of an appreciating effect of shocks to relative productivity levels. Real exchange rates' response to such shocks, when having the sign predicted by the BSH, tends to be short-lived. On the contrary, in a number of occasions we find that fiscal shocks, especially those of a considerable size, like the prolonged fiscal deficits in Italy in recent years and during the interwar period, do have persistent effects on the real exchange rate[36]. This is at odds with what generally expected, as one would

[36] Such effects are line with the predictions of the theoretical model adopted here in the first case, but not in the second, thus highlighting an important difference in the transmission of fiscal shocks.

imagine the effects of demand-side shocks to be wiped-out by factor mobility in the medium to long-run. In the next section we briefly summarise our findings, and point to possible avenues for future research.

## 6. Concluding Remarks

In this chapter, we aimed at providing some empirical evidence as to the origins of real exchange rate fluctuations over a long period of data. Time-varying VAR estimates, conducted within a Kalman filter perspective, in which stringent identifying restrictions need not to be imposed, sought to attribute those fluctuations to historical episodes of shocks to productivity levels and government spending. Overall, we found very weak evidence in favour of shocks of the first kind as having a persistent appreciating effect on the real exchange rate. Somehow, stronger support is found in favour of such an effect as produced by shocks to the relative fiscal position[37]. Aside from this general conclusion, three points seem to emerge from the empirical exercise we performed in the previous section.

First, estimation over sub-samples that do not include the war and inter-war periods shows mixed results as regards the effect of divergences in productivity levels on the real exchange rate. Especially during the Gold Standard years, such divergences tend to yield a real depreciation, whereas the case for a real appreciation, confirming the Balassa-Samuelson hypothesis, appears marginally stronger in the post-war and post-Bretton Woods periods. Moreover, we have found some evidence supporting the idea that during the Gold Standard fiscal imbalances were producing real depreciation of the exchange rate, exactly the opposite than during the post-war period.

Second, estimates over the entire sample show that, for particular rates and specific episodes, like the lira/dollar in very recent years, prolonged and sustained differences in government spending net of tax revenue appear to trigger persistent real appreciation. During the Gold Standard, and the early Bretton Woods years, however, the same shock seem leads to a diametrically opposite response of the real exchange rate. This can be readily interpreted as evidence of the fact that, beyond the exchange rate regime in place, which is indeed relevant, what matters in determining the persistence of shocks to

[37] This result is broadly in line with what found by Clarida and Gali (1994) for the post-Bretton Woods period.

the real exchange rate is the underlying economic structure. That is, there appears to be some evidence pointing to the underlying economic structure and the exchange rate regime as determining real exchange rate volatility, and the transmission of real shocks to the terms of trade.

It is then true, as we hinted, that the interactions between productivity levels, the fiscal balance and the real exchange rate be might strongly affected by the nature of the international arrangements[38] linking the monetary authorities. The sub-sample results we found, however, confirm that the exchange rate regime is not the unique determinant of the transmission of such interactions. Beyond the international monetary arrangements in place, one may wonder about the direction in which causality runs along such links, and along the interactions between the behavioural relationships in the economic system and policy's institutional settings.

In the first chapter of this thesis, we have found some evidence that suggests how, in the context of monetary policy conduct, institutional reforms probably tend to somehow follow, rather than precede, changes in the collective preferences or in the basic structure of an economic system. In the same vein, one cannot therefore exclude that the exchange rate regime is relevant in accounting for differences in the transmission of fiscal and productivity shocks to the real exchange rate. However, changes in the monetary regime should be seen as an effect, not a cause, of changes in the behavioural relationships in the economy. Demographic and technological factors are likely to account for changes in relative price flexibility even to a greater extent than the exchange regime in place. Moreover, the apparent correlation between the fiscal stance and the behaviour of productivity, suggests that it is mainly through such linkages that the above factors affect the transmission of shocks. Accordingly, if the response of the real exchange rate to various shocks differs over time, this is likely to be due to all those factors. The response itself may be endogenous to the set of conditions that led to the adoption of the specific exchange rate regime in place. It then follows that whether real shocks tend to prevail on monetary shocks as to the generation of observed real exchange rate fluctuations, is a highly case-specific matter.

Though blurred by the lack of more precise measures of productivity and government spending for the early part of our sample, the results just summarised somehow help bridging the gap between theoretical predictions and observed real

[38] Or, perhaps, by the absence of such arrangements, as under free exchange rate floating regimes

exchange rate behaviour. In particular, many of the long-run hypotheses some recent advances in intertemporal macroeconomics have recently developed, deserve further empirical investigation, perhaps along the historical lines we begun to tackle in this study. One possibility would be to extend our time-varying approach to a model explicitly allowing for discrete regime changes, so that the passages from Gold Standard, to Bretton Woods, and to the free-floating regime is directly accounted for by the statistical model. Another research priority would call for the construction of more refined indicators of the policy stance, especially for fiscal policy. As we have seen, the complex links between monetary and fiscal policy shocks often prevented exact identification of the underlying dynamics. In addition, the results of the existing literature, as we have seen, make us suspect that one major problem of these analyses is the identification of productivity measures that are not as strongly correlated to the fiscal stance as in our and other cases. However, we believe that our study has unveiled some important features of historical real exchange rate movements, and successfully addressed a number of relevant methodological issues.

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**Data Appendix**

Real GDP and government expenditure series were obtained by deflating the current price series using the GDP deflator.

The pre-1950 GDP, fiscal and price data were obtained from Flora (1983, 1987), *State Economy and Society in Western Europe, 1815-1975. Vols. I and II*, Campus: Frankfurt, and Mitchell (1992), *International Historical Statistics – Europe 1750-1988*, New York: Stockton Press. and Mitchell (1993), *International Historical Statistics – The Americas 1750-1988*. New York: Stockton Press.

The pre-1950 data on interest rates were obtained from Spinelli and Fratianni (1991) *Storia Monetaria d'Italia*, Milano: Mondadori; Friedman and Schwartz (1982) *Monetary Trends in the UK and the US* , Chicago: Chicago University Press; and Homer and Sylla (1995) *A History of Interest Rates*, New Brunswick: Rutgers University Press.

The data for the post-1950 period was obtained from standard sources (*IFS, WEFA*), and spliced together to obtain continuous series for the whole sample period.

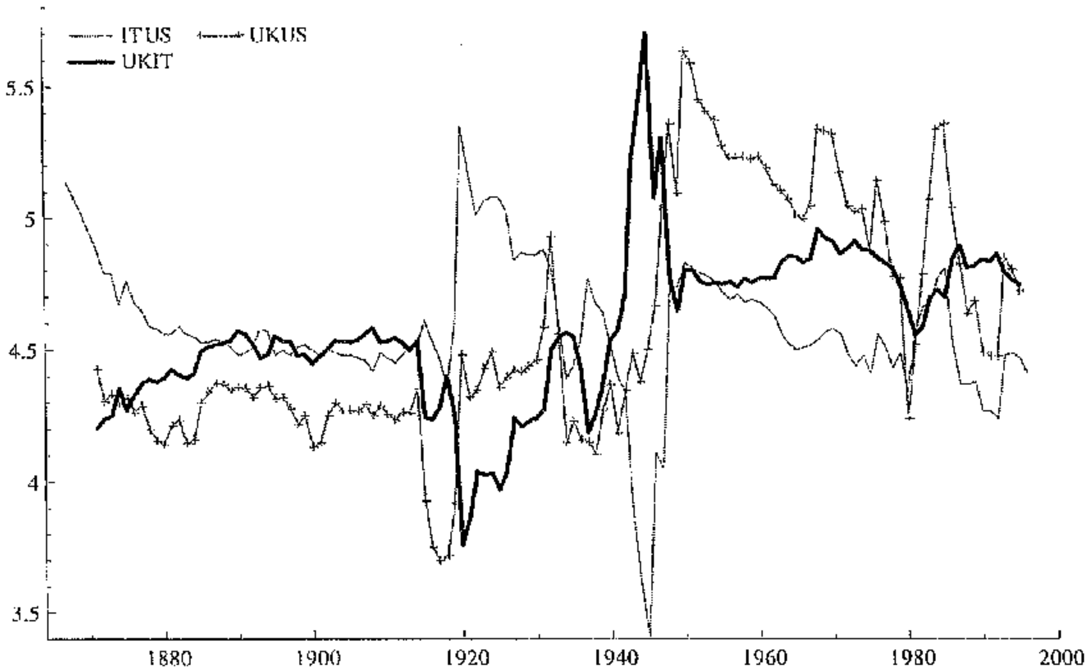



**Figure 1** – Bilateral real exchange rates (CPI), 1867-1996 (scaled means and ranges). UKUS, UK sterling/US dollar; UKIT, UK sterling/Italian lira; ITUS, Italian lira/US dollar.

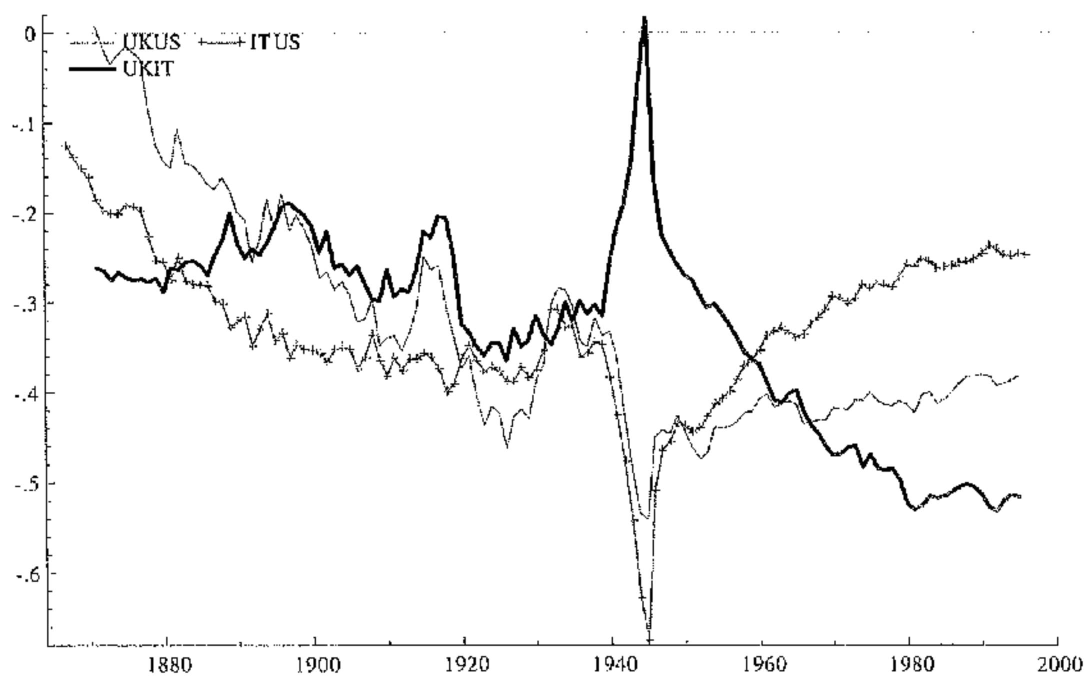



**Figure 2** – Differentials in productivity levels (see the main text for details on the proxies used here), 1867-1996 (scaled means and ranges).

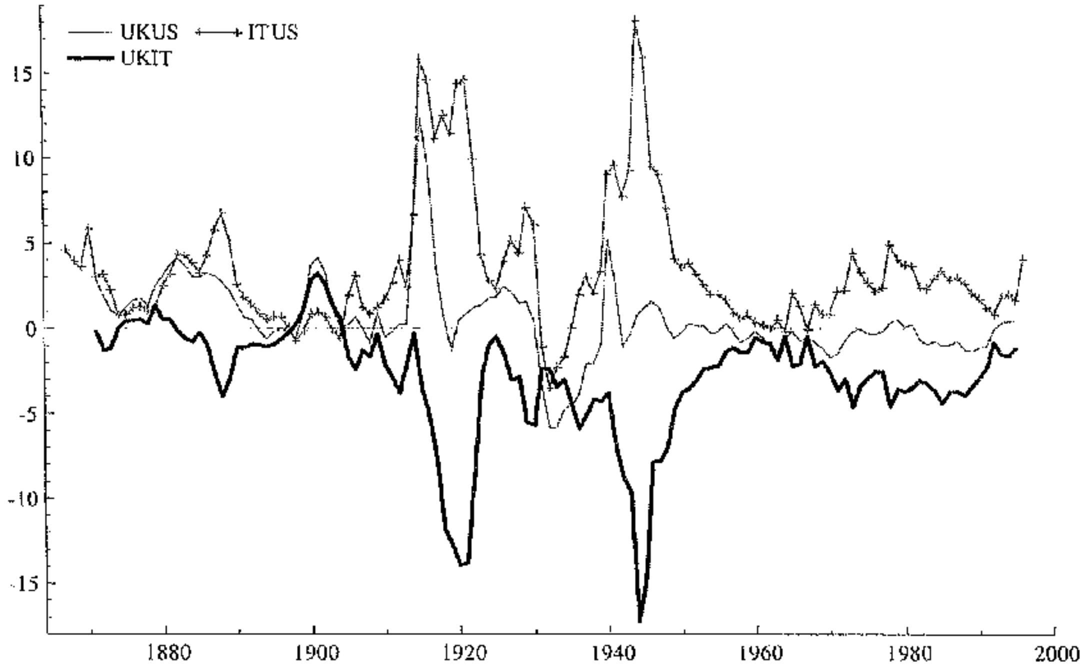



**Figure 3** – Differentials in budget deficit /real gdp ratios, 1867-1996 (scaled means and ranges).

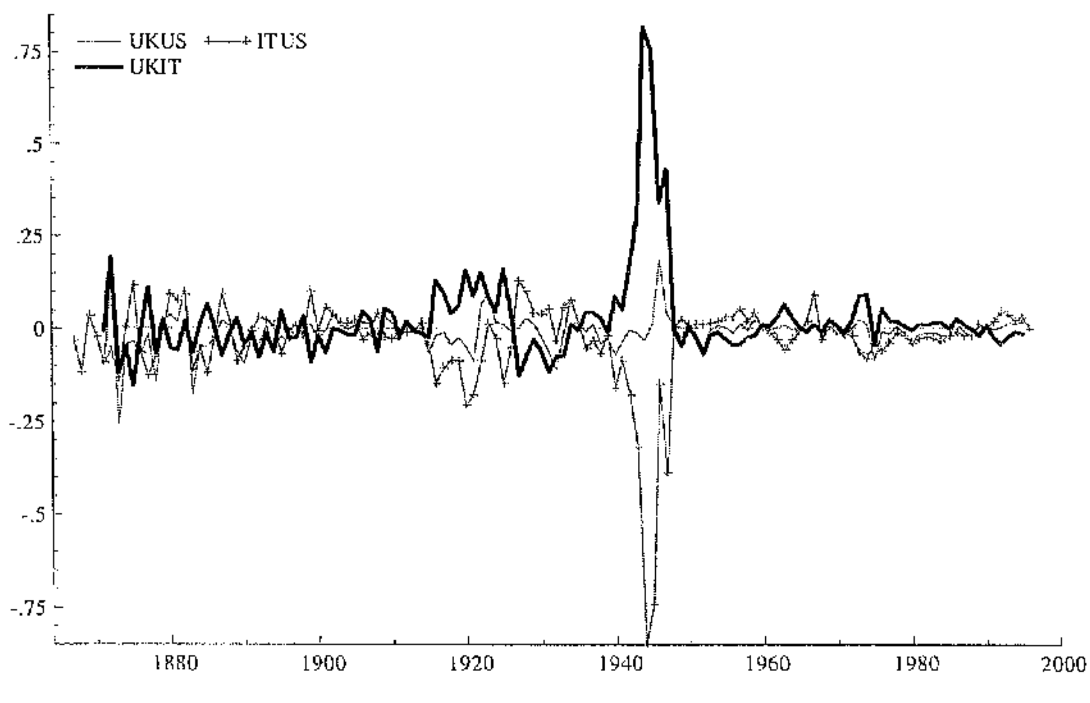



**Figure 4** – Real interest rate differentials, 1867-1996 (scaled means and ranges).

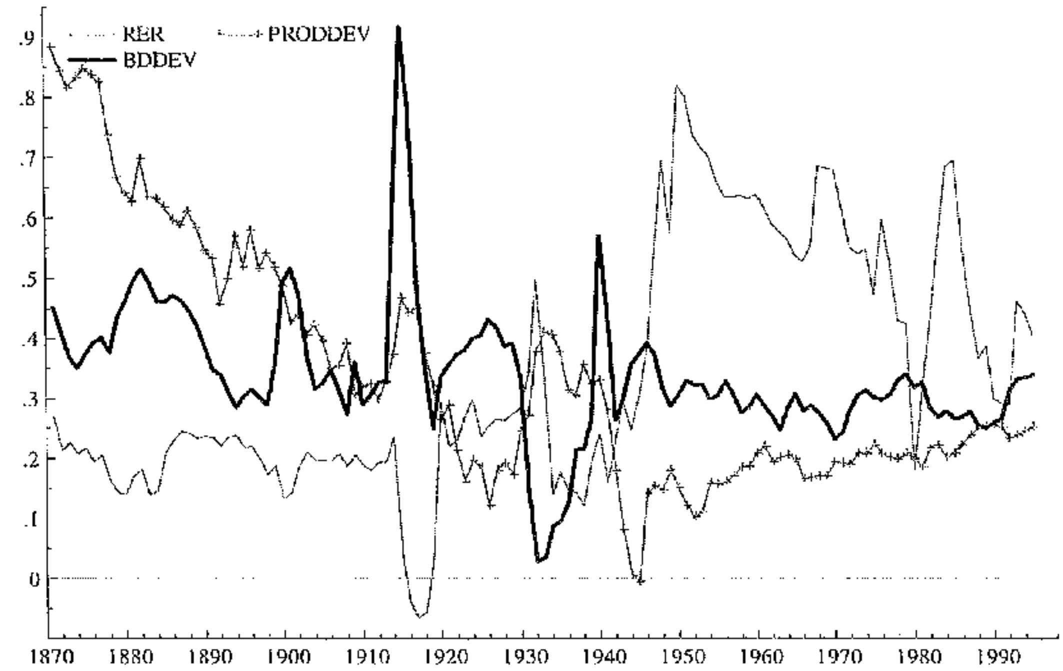



**Figure 5** – Real exchange rate, and differentials in productivity levels and budget deficit/GDP ratios. UK/US, 1871-1995 (scaled means and ranges)

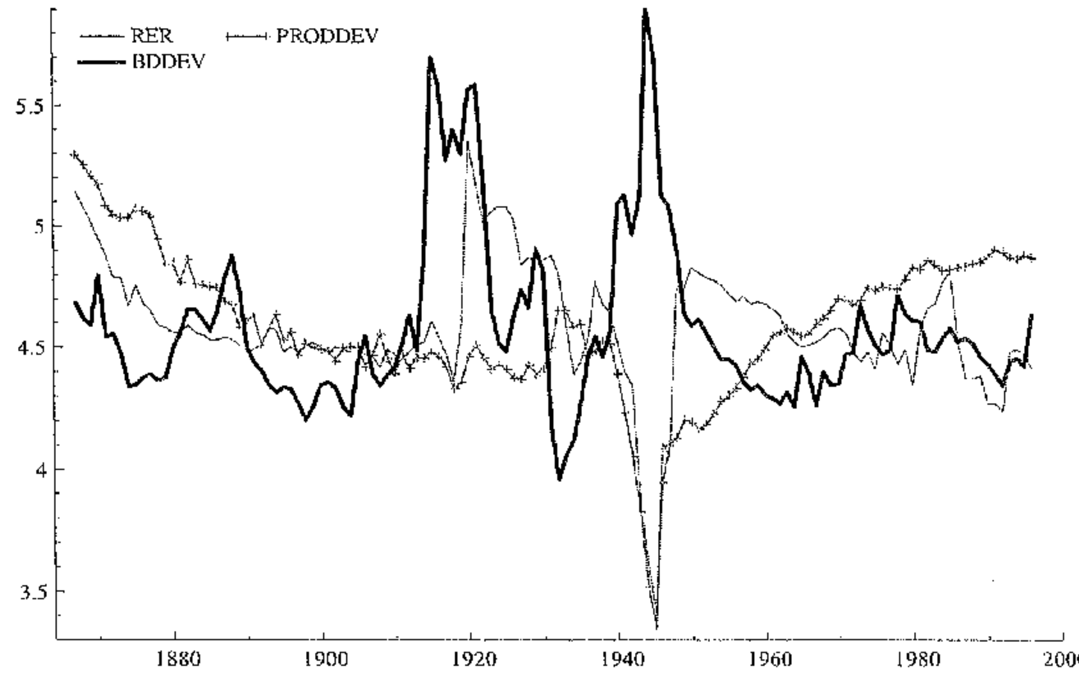



**Figure 6** – Real exchange rate, and differentials in productivity levels and budget deficit/GDP ratios. Italy/US, 1867-1996 (scaled means and ranges)

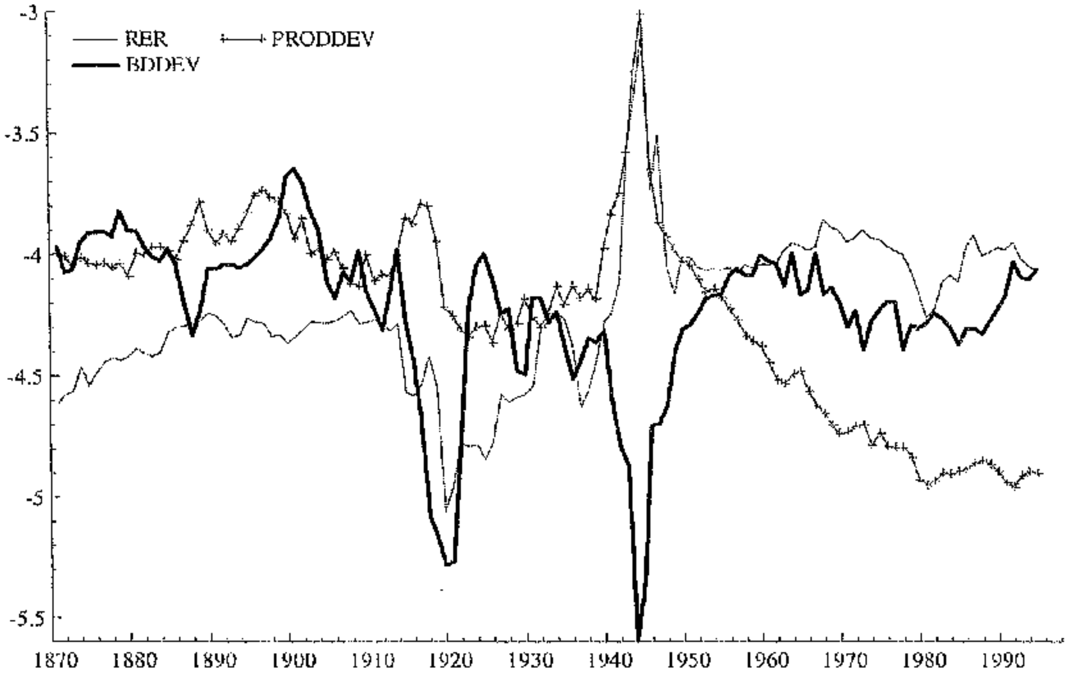



**Figure 7** – Real exchange rate, and differentials in productivity levels and budget deficit/GDP ratios. UK/Italy, 1871-1995 (scaled means and ranges)

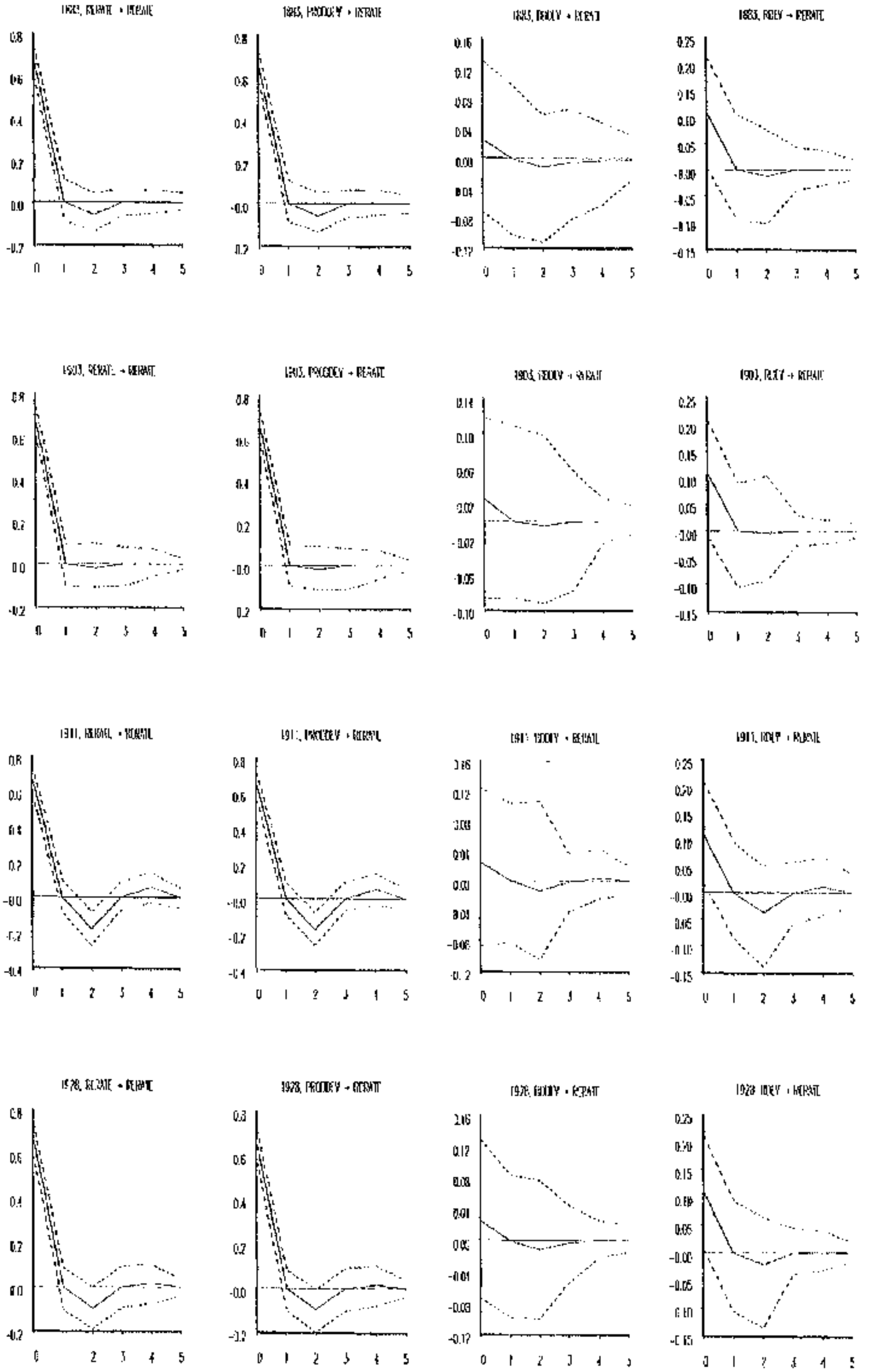


**Figure 8, Panel a** – UK sterling/US dollar, 1871-1995. Generalised Impulse Responses.

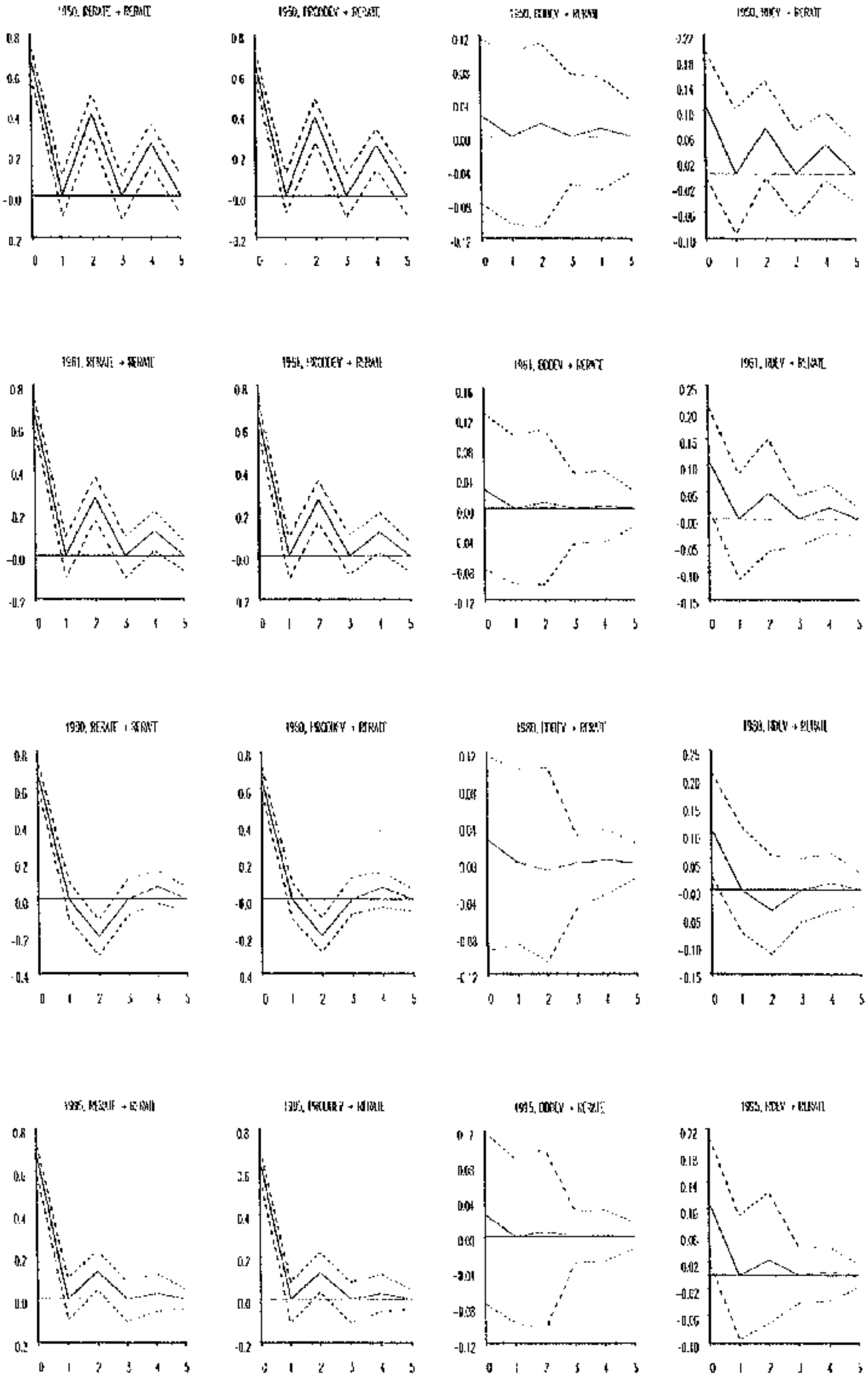


**Figure 8, Panel b –** UK sterling/US dollar, 1871-1995. Generalised Impulse Responses.

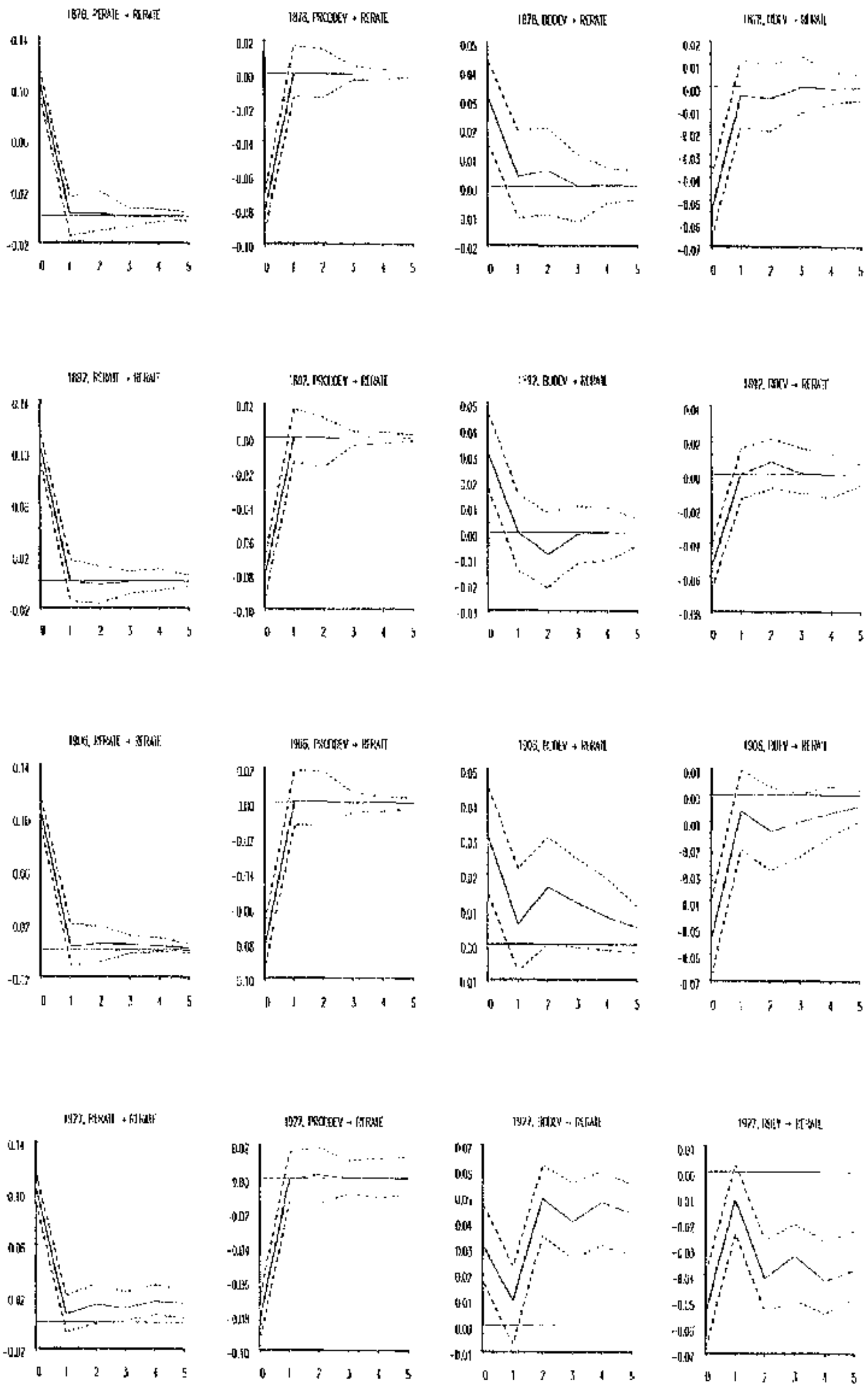


**Figure 9, Panel a** – Italian lira/US dollar, 1868-1996. Generalised Impulse Responses.

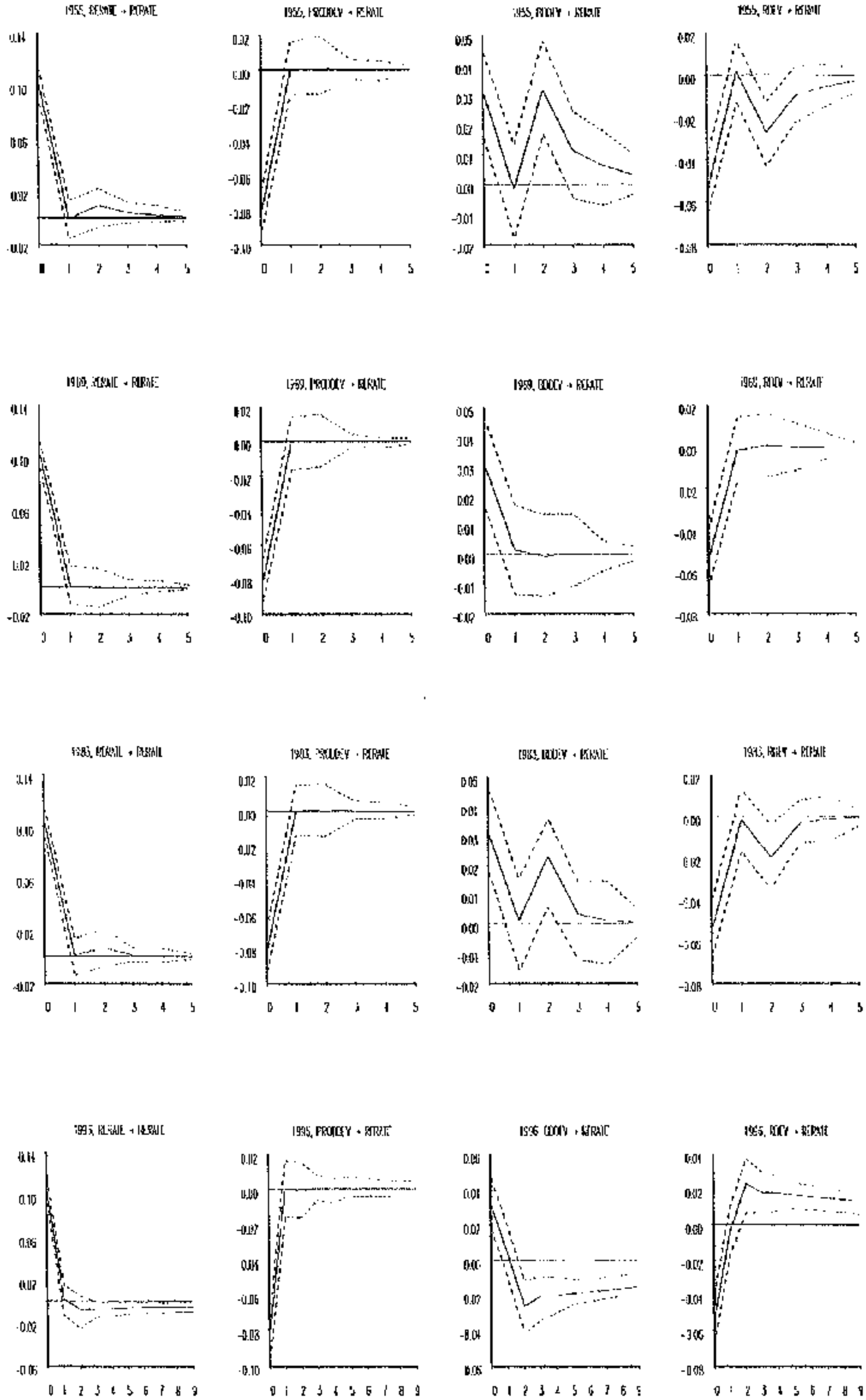


**Figure 9, Panel b** – Italian lira/US dollar, 1868-1996. Generalised Impulse Responses.

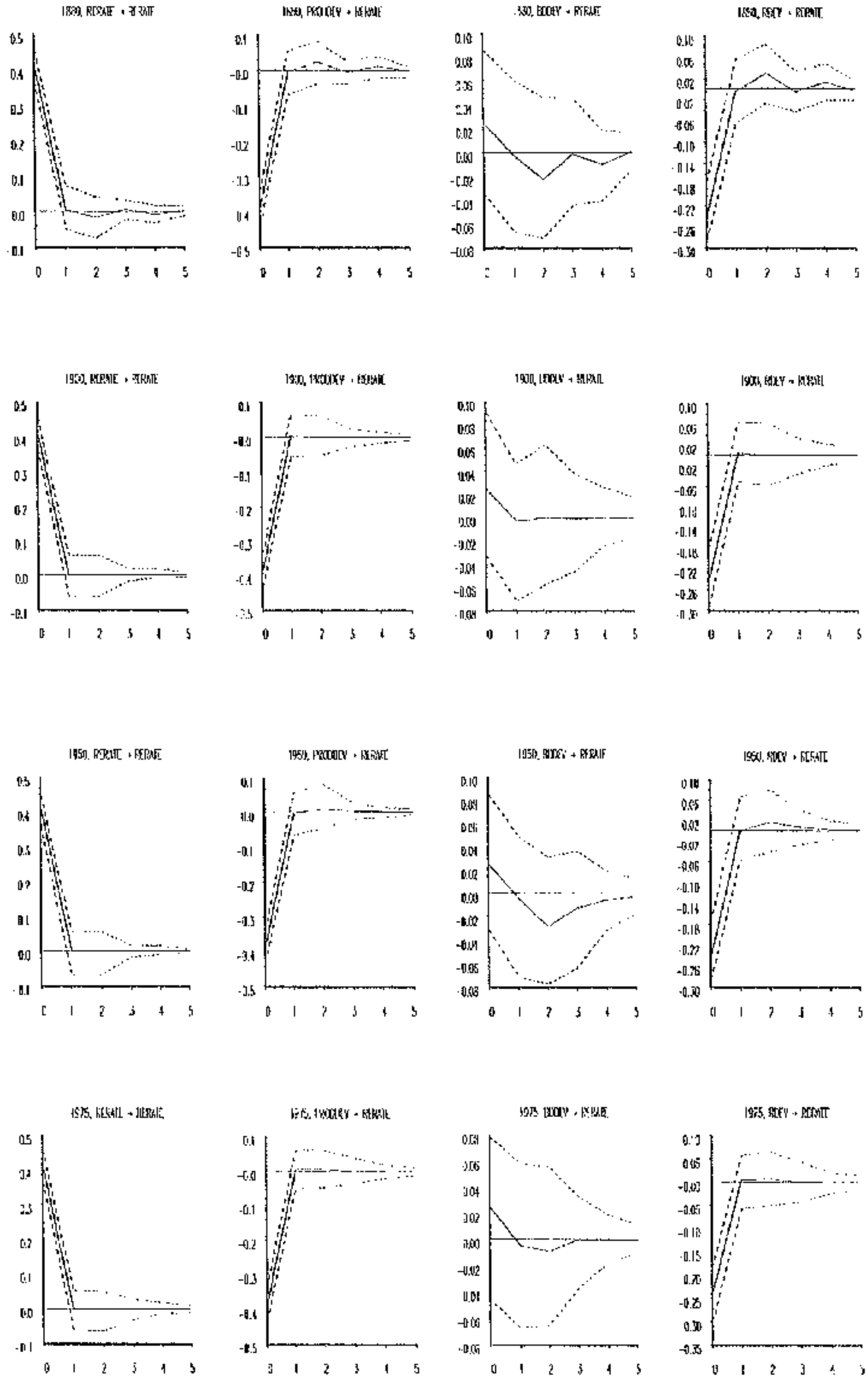


**Figure 10, Panel a** – UK sterling/Italian lira, 1871-1995. Generalised Impulse Responses.

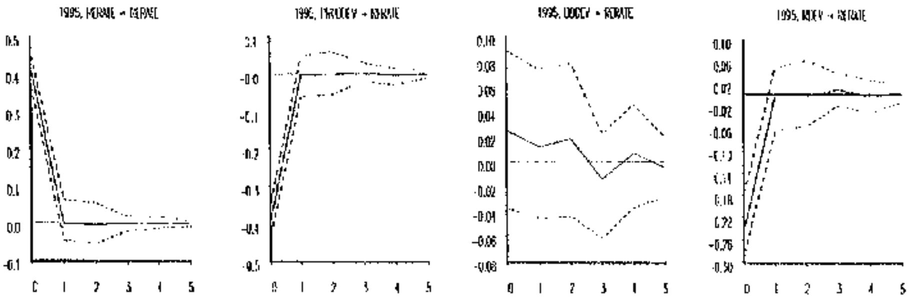
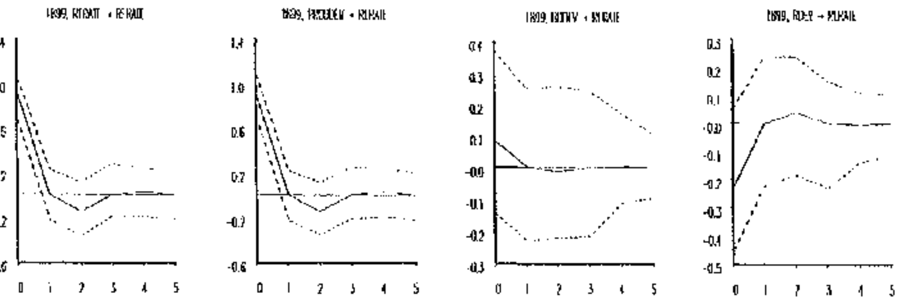
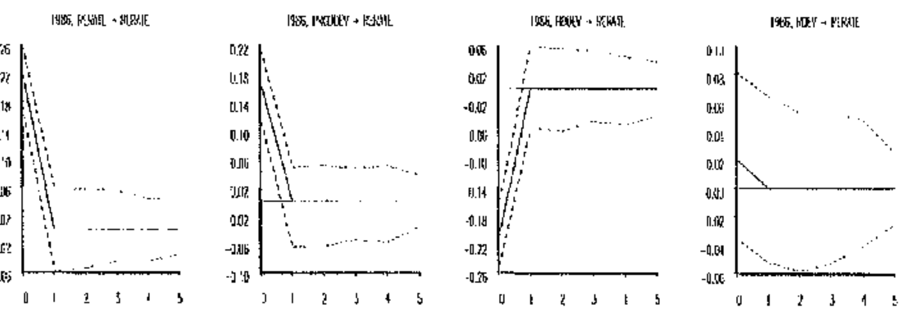


**Figure 10, Panel b** – UK sterling/Italian lira, 1871-1995. Generalised Impulse Responses.

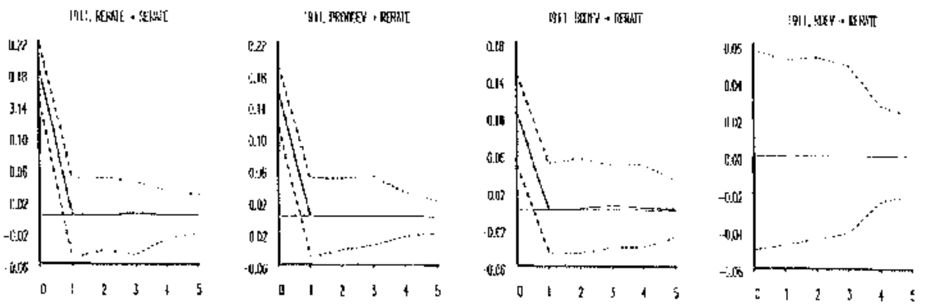


**Panel a** – UK sterling/US dollar, 1871-1913.

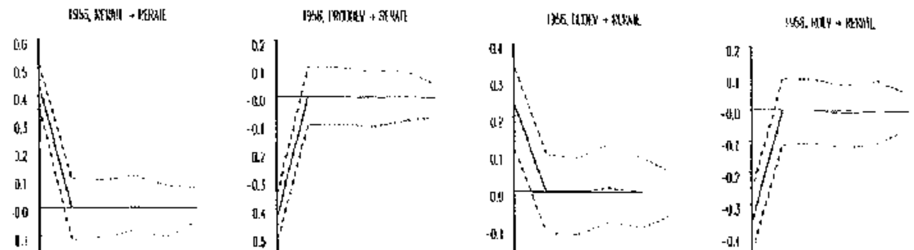


**Panel b** – UK sterling/US dollar, 1951-1995.

**Figure 11** - UK sterling/US dollar, various samples. Generalised Impulse Responses.

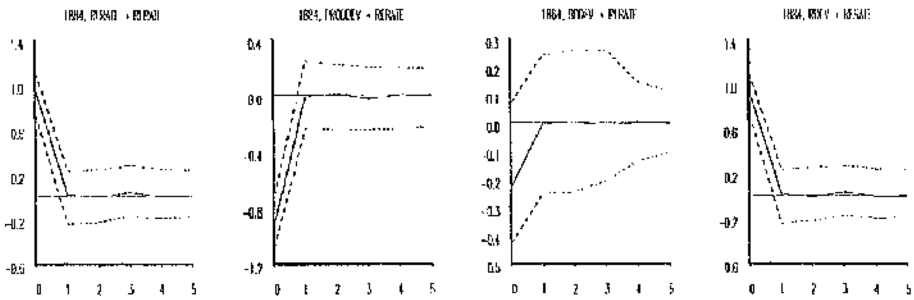


**Panel a** – Italian lira/US dollar, 1867-1913.

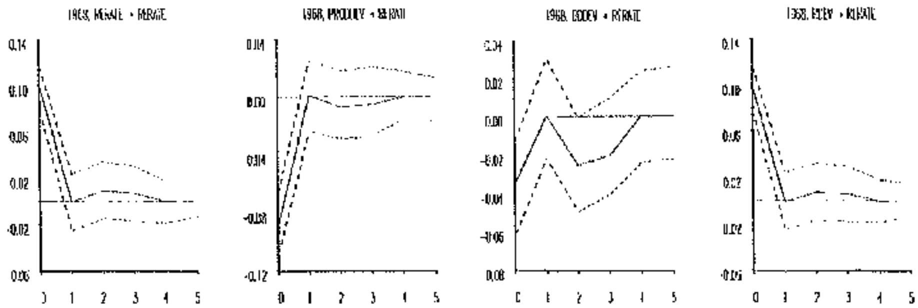


**Panel b** – Italian lira/US dollar, 1949-1996.

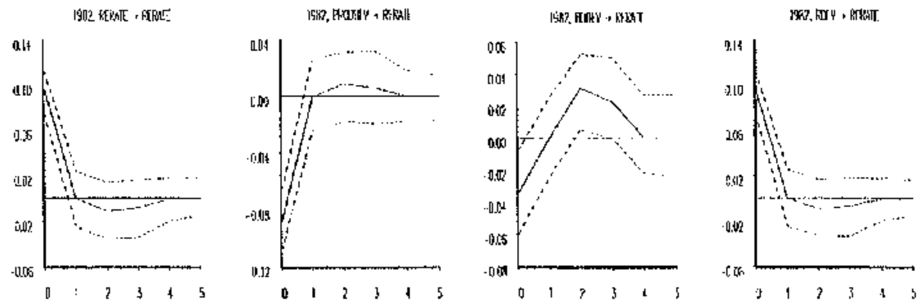
**Figure 12** – Italian lira/US dollar, various samples. Generalised Impulse Responses.



**Panel a** – UK sterling/Italian lira, 1871-1913.



**Panel b** – UK sterling/Italian lira, 1949-1995.



**Panel c** - UK sterling/Italian lira, 1949-1995.

**Figure 13** – UK sterling/Italian lira, various samples. Generalised Impulse Responses

## CONCLUSIONS

### Summary of Findings for Each Chapter

The aim of this thesis was to bridge the substantial gap existing between some of the theoretical predictions elaborated by the literature on the political economy of monetary policymaking, and the apparent lack of empirical validation which has often accompanied their emergence as normative guidelines for policymakers. Part of this research was further motivated by the need to establish some empirical regularity as to the role of leading indicators for monetary policy, and of fiscal and productivity shocks for the behaviour of real exchange rates.

Overall, this empirical study has addressed a number of relevant methodological issues, and unveiled some important features concerning the role of structural and institutional changes on contemporary policymaking. What follows is a summary of the main results, articulated for each chapter.

*Chapter One: Interest Rate Rules and Policy Shifts in OECD Economies*

In this chapter, we have discussed and estimated forward-looking interest rate reaction functions for two groups of OECD economies. Our aim was twofold. First, we sought to envisage whether the recent emphasis placed by the existing empirical literature on the consistency of monetary regimes in the G-3 economies with an inflation-forecasting approach, was validated by the data. Second, we wanted to establish whether there was any systematic pattern between institutional change, in the form of the adoption of explicit inflation targets and central bank reforms, and the operational conduct of monetary strategies. A number of general conclusions emerge from our empirical results.

First, the evidence we gathered suggests that the monetary authorities in our sample did not follow stable simple forward-looking policy reaction functions based on output gaps and expected inflation (and, *a fortiori*, Taylor rules). In the US and Japan, countries where there have been no major institutional reforms, we find that policies underwent considerable evolutions in the 1980s and 1990s. However, it is only since the 1990s that estimated interest rate rules in these countries begin to look like the ones the theoretical research on inflation-forecast targeting has recently illustrated.

Second, in countries where there were explicit intermediate targets (such as the growth of M3 in Germany), these appear mainly as a device to anchor expectations. In addition, monetary policy is often found to follow a broad set of external objectives. Our results confirm those of previous researches who have detected in the Bundesbank a marked "targeting" attitude regarding inflation, output, and some external conditions.

Third, with the exception of the UK, the recent switch to inflation targets in the countries we studied does not seem to have radically changed the way in which interest rate policy pursues its ultimate objectives. In practice, there is some evidence, particularly clear in Canada's and New Zealand's cases, that any major change in the responsiveness of interest rates to expected inflation took place well before the adoption of inflation targets. The same pattern seems to have been followed even when such institutional reforms have been accompanied by greater central bank independence. A possible interpretation is that the new regimes were brought in simply to consolidate gains in terms of lower inflation.

Finally, we detected some important differences in the behaviour of central banks as far as output stabilisation is concerned. On the one hand, we found some evidence suggesting that the Fed has been apparently exploiting its consolidated antinflationary reputation to focus on the cycle. At the other extreme, some monetary authorities apparently feel the need to build up such a reputation. .

Whether central banks that have only recently acquired their independence, such as the Bank of England and the European Central Bank, will find themselves into the first or the second category of experiences, remains, however, a fairly open question.

## *Chapter Two: Interest Rate Rules and Policy Credibility on the Road to EMU*

Similarly to what we did in the previous chapter for other countries, we have examined the behaviour of interest rate policies in France, Italy, Belgium and Ireland, during the eighties and on their road to EMU. Estimated reaction functions for each country provided us with some evidence about the relative costs these countries faced in maintaining their exchange rate parities within the ERM, and in fulfilling the basic convergence criteria laid out in the Maastricht Treaty. The role of exchange rate risk, and the way in which long-term interest rate differentials reflected the overall credibility of fiscal policies in each country, has been analysed in relation to the trade-offs national policymakers faced between domestic and external objectives.

Our findings highlight quite clearly that in some cases (France, Belgium) the process of convergence towards a stable exchange rate vis-à-vis Germany and a tight inflation control took place amidst sustained policy reforms over a medium- to long-term horizon. In addition, the ERM turbulence in 1992-93 does not appear to have impaired the largely achieved macroeconomic stability. On the contrary, Italy's, and to a much lesser extent, Ireland's monetary policies, seem to have been severely constrained, in their efforts to stabilise the economy, by the lack of credibility of their respective fiscal stances.

Overall, our results appear as a first empirical validation, with reference to European countries, of the idea that fiscal imbalances do impose some additional constraint on the manoeuvrability of monetary policies. This message appears even more relevant in the context of a unified European monetary policy process.

In the EMS context, the relationships between a country's budgetary position and interest rate determination were somehow made more evident by the presence of a further binding constraint on economic policies –the ERM band. The adoption of exchange rate agreements made financial markets, and the public in general, more aware of the new trade-off faced by monetary authorities, by attaching extra penalties to departures of interest rate policies from an anti-inflationary stance. It is clear that financial markets and policy observers, for the very same reason, scrupulously and critically monitor the behaviour of the European Central Bank. Whether the interesting results found in this study are to be observed for the euro area as a whole and for other economies, is left as a direction for further research.

*Chapter Three: Evaluating the Information Content of Euro Area Monetary Aggregates*

In this chapter, we provided some preliminary empirical evidence as to the usefulness of the growth of M3 for predicting future inflation in the euro area. First, we found little empirical support for rejecting at standard confidence levels Granger non-causality of M3 on prices. This conclusion is found stable throughout the sample and robust to a number of robustness checks. Second, we investigated the leading indicator properties of M3, by looking at a *Pstar*-type model in which information about the cyclical state of the economy and a measure of authorities' inflation target feed back onto the generation of inflation forecasts. Our results confirm that a significant positive association exists between the real money gap and future inflation up to five to six quarters ahead. Similar predictive ability is displayed by the output gap, the real interest rate and (to a lesser extent) by the term spread.

Third, we compared the forecasting properties of the model developed here to those of an alternative, non-monetary model for euro-area GDP. Overall, each model appears to have some strengths of its own: both of them incorporate some information that is relevant to explain GDP inflation. However, taken individually, none of the models seems to be able to provide a complete account of inflation developments in the euro area.

All this suggests quite clearly that information derived from both monetary and non-monetary models (and indicators) should be used to generate reliable forecasts for euro-area inflation.

*Chapter Fourth: Real Exchange Rates, Productivity Levels and Government Spending*

In the final chapter, we aimed at providing some empirical evidence as to the origins of real exchange rate fluctuations over long historical datasets for the US, UK, and Italy. Time-varying VAR estimates, conducted within a Kalman filter perspective, in which stringent identifying restrictions need not to be imposed, sought to attribute those fluctuations to historical episodes of shocks to productivity levels and government spending.

Our main result is that there appears to be very weak evidence in favour of shocks of the first kind as having a persistent appreciating effect on the real exchange rate. Somehow, stronger support is instead found in favour of such an effect as produced by shocks to the relative fiscal position. Additional points emerge from the empirical exercise we performed.

First, estimation over sub-samples that do not include the war and inter-war periods shows mixed results as regards the effect of divergences in productivity levels on the real exchange rate. Especially during the Gold Standard years, such divergences tend to trigger a real depreciation, whereas the case for a real appreciation, confirming the Balassa-Samuelson hypothesis, appears marginally stronger in the post-war and post-Bretton Woods periods. Moreover, we have found some evidence supporting the idea that, during the Gold Standard, fiscal imbalances were producing real depreciation of the exchange rate, exactly the opposite than during the post-war period.

Second, estimates over the entire sample show that, for particular rates and specific episodes, like the lira/dollar in very recent years, prolonged and sustained differences in government spending trigger persistent real appreciation. During the Gold Standard, and the early Bretton Woods years, however, the same shock leads to a diametrically opposite response of the real exchange rate. That is, there appears to be some evidence pointing to the underlying economic structure and the exchange rate regime as determining real exchange rate volatility, and the transmission of real shocks to the terms of trade. Moreover, the sub-sample results we found confirm that the exchange rate regime is not the unique determinant of the transmission of such interactions.

In the first chapter of this thesis, we have found some evidence suggesting how, in the context of monetary policy conduct, institutional reforms probably tend to follow, rather than precede, changes in collective preferences or in the basic structure of an economic system. In the same vein, one cannot exclude that the exchange rate regime is relevant in accounting for differences in the transmission of fiscal and productivity shocks to the real exchange rate. However, changes in the monetary regime should be seen as an effect, not a cause, of changes in the behavioural relationships in the economy. Demographic and technological factors are likely to account for changes in relative price flexibility even to a greater extent than the exchange regime in place. Moreover, the evident correlations between fiscal stance and productivity, suggests that it is mainly through such still unclear linkages that the above factors affect the transmission of shocks. The transmission mechanism itself may be endogenous to the set of conditions that led to the adoption of the particular exchange rate regimes in place.