MONETARY POLICY REGIME SHIFTS: 
NEW EVIDENCE FROM TIME-VARYING 
INTEREST RATE RULES 

by 
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Discussion Paper n. 0602
Monetary Policy Regime Shifts: New Evidence from Time-Varying Interest-Rate Rules

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January 2006

Abstract

We estimate forward-looking interest-rate rules, for major advanced countries, allowing for time variation in their parameters. Traditional constant-parameter reaction functions likely blur the impact of i) model uncertainty, ii) conflicting objectives, iii) shifting preferences and iv) nonlinearities of policymakers choices. We find that monetary policies followed by the US, the UK, Germany, France and Italy, often described in terms of standard Taylor rules, are best summarized by feedback rules that allow for time variation in their parameters. Estimated rules point to sizeable differences in the actual conduct of monetary policies, even in the countries now belonging to the EMU. Also, our TVP specification outperforms the conventional Taylor rule in tracking the actual Fed funds rate.

JEL Codes: E52, E58, E60

1 Introduction

Feedback rules nowadays are customary tools for monetary policy analysis. The latest generation of quantitative models of the monetary transmission mechanism supports a wide range of stabilization policies (see Woodford, 2003). On the other hand, rule-based policymaking is generally accepted as the framework for monetary policies strategies (see ECB, 2004; Bank of England, 1999).

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Interest-rate policies are complex decisions: they rely on a multitude of indicators and models, and they are by nature associated to events not always captured by relatively simple econometric approaches. Nonetheless, there is now a wide consensus that the class of simple policy rules examined by recent microfounded models of the economy generates stabilisation properties that are very close to those of optimal feedback rules (again, see Woodford, 2003).

In practice, monetary policy reaction functions reflect the importance that policymakers attach to often conflicting objectives, as well as their views on the structure of the economy. As such, they might display instability and parameter variation. From this point of view, traditional constant-parameter reaction functions are likely to blur the impact of i) model uncertainty, ii) conflicting objectives, iii) shifting preferences on policymakers’ choices, and iv) nonlinearities.

In particular, intuition suggests that models for policy analysis that do not allow for shifts in behavioural relationships can produce misleading results. Dynamic stochastic general equilibrium (DSGE) models represent one potentially promising way of overcoming these problems. This approach is increasingly popular as it is based on models derived from microfounded descriptions of the economy, often very rich in terms of their ability to capture all interactions between economic agents. Recent studies also provide relatively simple techniques to estimate and evaluate linearized DSGE models using Bayesian methods (see Del Negro and Schorfheide, 2004; Del Negro et al., 2005). A Bayesian framework is a pretty natural setup to account for model and parameter uncertainty. However, the DSGE approach imposes a large number of restrictions on the data, and in the context of interest-rate rules its usage to make qualitative assessments on monetary policy conduct appears problematic. However, a single-equation approach to policy rules, like those proposed by Clarida et al. (1998, 2000) or, more recently, by Rudebusch (2005), lend itself to estimation in a Bayesian framework. This paper attempts to provide some evidence on interest-rate policies in major economies through the implementation of a loosely-defined Bayesian approach to the estimation of its coefficients.

We estimate policy rules as summarized by simple reaction functions defined in terms of expected inflation and output gaps. In contrast with most existing analyses, we employ a time-varying-parameter (TVP henceforth) approach based on the Kalman filter algorithm to estimate our instrument rules. In practice, we allow the policy rules’ coefficients to vary over time. What we obtain are estimates of the state vector for each observation in our sample. These estimates describe the evolution of monetary policy over
A host of factors may cause shifts in the parameters of estimated monetary policy reaction functions. Some of them may be due to institutional reforms, like the introduction of inflation targets or the move to a different monetary policy framework, like an exchange-rate peg or a fixed exchange rate. In estimating reaction functions defined in terms of final policy objectives, one aims at capturing the relative emphasis placed by policymakers to the attainment of output and inflation targets. Over time, such emphasis might change. The approach followed to capture these shifts should somehow reflect the observer’s view on the likelihood that they are smooth transitions or abrupt regime changes. As better argued in the next Section, our description of actual policies is better suited than conventional fixed-parameter studies in dealing with Lucas critique, and more in general with the issue of structural change. Moreover, by permitting the weights of the policy rule to vary gradually, we are able to identify relevant shifts in policy conduct, regardless of their exact configuration.

We find that monetary policies followed by the US, the UK, Germany, France and Italy, often described in terms of standard Taylor rules, are best summarized by feedback rules that allow for time variation in their parameters. Overall, the evolution of coefficients portrays a richer picture of central banks’ conduct, and it is consistent with historical macroeconomic events. Moreover, estimated policy rules summarize policies according to different coefficient estimates, and, in some cases, different functional forms. Estimated parameters tend to shift over time, in most cases in a smooth and gradual fashion. These findings corroborate our choice of estimating single-country interest-rate rules, and the validity of a TVP approach to this issue. Finally, our rule appears to dominate the conventional specification of the Taylor rule in tracking interest-rate developments.

The paper proceeds as follows. Next Section discusses some recent developments in the literature on monetary policy and structural change, which motivate our approach. Section 3 describes our model of simple interest-rate rule, while Section 4 presents our estimation results. Section 5 concludes.
2 Interest-Rate Policy: Optimal and Estimated Rules

2.1 A Reference Model

Generally speaking, inflation (-forecast) targeting rules, i.e., rules that describe policy rates as following the projections of inflation and some additional variables in the future, have become the centerpiece of modern monetary policymaking. These rules are usually derived from models that posit optimization and rational expectations. More in detail, these models often assume forward-looking pricing behaviour and sticky prices\(^1\).

A baseline version of such models often comprises an intertemporal IS relation,\(^2\)
\[
y_t = E_t y_{t+1} - \gamma E_t (i_t - \pi_{t+1} - r^n_t) \tag{1}
\]
and an aggregate-supply relation,\(^2\)
\[
\pi_t = \delta y_t + \beta E_t \pi_{t+1} + \eta_t \tag{2}
\]
where \(y_t\) is the output gap, \(i_t\) is the nominal short-term (policy) interest rate, and \(\pi_t\) is inflation. Equation (2) is a structural Phillips-curve relation in which the link between current inflation and output (or the output gap) depends on the particular price-setting mechanism (in Calvo-type models, by staggered pricing). Equation (1) is an approximation of the log-linearized version of a consumption Euler equation. \(r^n_t\) and \(\eta_t\) are two exogenous disturbances, interpretable, for instance, as a) a shock to the equilibrium real rate of interest (see Woodford, 2003), and b) a cost-push shock, respectively\(^2\).

The above relations constrain stabilization policy. The model assumes that authorities aim at minimizing the expected value of an intertemporal loss criterion, whose period loss function can be
\[
L_t = \pi_t^2 + \lambda_y (y_t - y^*)^2 + \lambda_i (i_t - i^*)^2 \tag{3}
\]
\(y^*\) and \(i^*\) are target values for output and interest rates. Without loss of generality, we assume zero as the target for inflation. Therefore, drawing on the extensive literature on monetary policy rules, we describe the behaviour

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1 See Woodford (2003) for many examples. We follow his small quantitative model in the discussion below.
2 This class of models do not refer universally to a neo-Wicksellian framework. However, the explicitly New Keynesian paradigm of models like the one just described is widely accepted.
of the nominal policy rate according to the following implicit instrument rule3:

\[ i_t = (1 - \varphi_1) i^* + \varphi_1 i_{t-1} + \varphi_2 \Delta i_{t-1} + \varphi_3 \pi_t + \varphi_4 \Delta y_t \]  

(4)

where the \( \varphi \) coefficients are combinations of the various structural parameters of the model. According to this rule, the policy instrument displays persistence4, follows developments in inflation and the output gap, and responds to recent interest-rate changes.

It must be said that the above framework is just a sketch-like description of the variety of available models. Also, one can certainly add a number of salient features to this specification. For instance, given the observed inertia in inflation behaviour, models with partial indexation to an aggregate price index or staggered wage and price contracts generate persistence in inflation (Christiano et al., 2005). The form of the optimal rule itself changes somewhat when both prices and wages are sticky (Erceg et al., 2000; Amato and Laubach, 2003), or if private expenditure displays habit persistence (as in Smets and Wouters, 2003, or Muscatelli et al., 2004).

Finally, in the real world we observe well-documented lags in the reaction of output and inflation to unexpected changes in interest rates. Many authors agree that to match this regularity one should anyway allow pricing and spending decisions to be predetermined. This also motivates the minimization of interest-rate volatility as an additional goal of stabilization policies in equation (3).

2.2 An Overview of Applied Literature

This intense theoretical literature has spurred many empirical estimates of simple monetary policy rules5. Since Taylor’s (1993) proposal of a simple interest-rate rule defined in terms of current inflation and output gap6, central bank behaviour has been studied through estimated versions of equation (4) on quarterly or monthly data. Examples include Batini and Haldane (1999), Clarida, Galì and Gertler (2000), and Muscatelli et al. (2002b).

One important issue has partly remained in the background of this burgeoning literature: how does Lucas critique affect the evaluation of monetary policy models? Available estimated policy rules acknowledge that monetary

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3 The specification below is pretty standard in applications based on quarterly data. See for instance the studies in Taylor (1999).

4 As suggested, inter alia, by Rotemberg and Woodford (1999).

5 Many studies included additional variables to a baseline specification in terms of inflation and output to account for intermediate or alternative central bank targets.

6 See Taylor (1999) for an estimated version of this proposal.
policymakers have displayed changing attitudes towards inflation and output stabilization. For instance, Taylor (1999) and Clarida et al. (2000) find different parameter estimates over sub samples of the last three decades of US data. Muscatelli et al. (2002b) identify gradual changes (and some regime shifts) in the conduct of interest-rate policies of major advanced countries.

According to Lucas critique, these changes could alter agents’ expectations of the future policy course, and, in our context, modify the behavioural parameters of equations (1) and (2). Forecasting and policy analysis conducted over models whose behavioural equations react to structural policy changes could yield, in principle, to biased results. However, Rudebusch (2005) points out that, when one conducts stability analysis on reduced-form representations of the economy, like the VAR models so popular in the literature on the transmission mechanism, results often support structural invariance. Estrella and Fuhrer (2003) and Rudebusch (2005) argue that while relevant at a theoretical level, Lucas critique might be less so on an empirical ground, especially if investigated with reference to Fed’s reaction function. In other words, Lucas critique would not apply in the context of small shifts in the policy rule. This argument introduces some realism in the empirical investigation of policy regimes. However, it also calls for further caution, since it implies that some policy shifts might be detectable through econometric methods despite being irrelevant as to its effects on agents’ expectations.

One possibility is that changes in policymakers’ preferences are small and gradual, so that they are processed by private agents through a similarly gradual learning process, difficult to capture by overparameterized, constant-parameters representations of the economy as VAR models typically are. For instance, many studies, starting from Clarida et al. (1998) simply assume that US monetary policy has undergone a single break, often corresponding with the start of Chairman Volcker’s tenure. Alternatively, there is a widespread conviction that monetary policies in the 1970s were characterised by a different vigour as to inflation control relative to those followed in the 1980-90s. Some authors (notably Taylor, 1999) conduct sub-sample estimation of policy reaction functions, finding significant differences in estimated parameters. Sims and Zha (2004) estimate a number of mul-

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7 The start of tenures of Paul Volcker (1979Q3) and Alan Greenspan (1987Q2) as Chairman of the Fed, as well as the end of the "monetarist experiment" engineered during Volcker’s chainmanship (1982Q3), are often held as break dates for US monetary policy.

8 For an exception see Muscatelli and Trecroci (2002).

9 Estrella and Fuhrer (2003) is another example.

10 Since Taylor (1999), the literature has stressed that until 1979 Fed’s policies violated
multivariate models with discrete breaks, and allow for simultaneity and regime switching in coefficients and variances. Among models with changes in equation coefficients, the best fit is found for a model that allows coefficients to change only in the monetary policy rule.

In contrast, Primiceri (2005) models and estimates a structural VAR of the US economy, allowing for time variation in both the coefficients and the covariance matrix of the model’s innovations. His estimates support the view that important changes affected both the systematic and nonsystematic components of US monetary policy over the past 40 years. In contrast with Sims and Zha (2004), the model studied in Primiceri (2005) captures changes in private sector behaviour. In the context of VAR models of the economy this is a particularly useful feature. The equations describing the private sector are convolutions of the underlying behavioural relationships, and for this reason one cannot clearly disentangle policy and non-policy responses.

Assenmacher-Wesche (2006) conducts an interesting analysis on monetary policies in the biggest EU countries and in the US. She estimates a Markov-switching model that allows for independent switching processes in the parameters of policy rules and in their residual variance. The results show that the inflation and output coefficients of the policy rule do evolve according to two different regimes. In particular, the evidence supports the existence of an "hawkish" regime (low coefficient on output stabilisation and a high one on inflation stabilisation), and a "dovish" regime (where the opposite holds). Moreover, one-off shocks like supply-side disturbances or institutional reforms are accounted for by the switching in the residual variance. The focus of Assenmacher-Wesche (2006) on changes in central banks’ attitude towards stabilisation targets makes it very close in spirit to our paper.

The problem with discrete-break (and conventional VAR) models is that they fail to account for gradual policy changes and lead to problematic interpretations in the case of regime shifts that do not fit properly into one of the modelled regimes. For instance, it would be interesting to extend the approach of Assenmacher-Wesche (2006) beyond the stylized two regimes allowed in the study\textsuperscript{11}. One promising route to account for these and other problems is the DSGE-VAR approach proposed, inter alia, in Del Negro et al. (2005). However, DSGE models are inappropriate tools for the evaluation of simple policy rules, given their interpretative complexity.

\textsuperscript{11} Another limitation in Wesche (2003) is that the degree of interest-rate smoothing is not allowed to switch across different regimes.

the so-called "Taylor principle", that is, they failed to raise nominal interest rates by more than the amount by which inflation exceeds the target.
Stability analysis on estimated forward-looking interest-rate rules provides some indications about changes that take place in the actual conduct of monetary policies. Indeed, Muscatelli et al. (2002b) find evidence that policy shifts in advanced countries’ monetary policies did take place, possibly reflecting shifts in collective preferences towards the relative costs of inflation. Even in countries, like the US and Japan, where monetary policies were unaffected by major institutional innovations as the introduction of inflation targeting, policy shifts were clearly visible. These findings motivate an investigation of policies focusing on gradual policy change rather than on structural breaks.

The policies of countries now part of the European Monetary Union have received remarkable attention in the applied literature on interest-rate rules. Most studies focus on single-country reaction functions (see Clarida et al., 1998; Muscatelli et al., 2002a), while others propose the estimation of area-wide policy rules. In the latter strand, Gerlach and Schnabel (2000) is one of the most influential contributions. These authors find that a modified version of the Taylor rule, specified in terms of area-wide inflation and output gaps, explains well the behaviour of average interest rates in the EMU countries in 1990-1998, with the exception of those years (1992-93) that saw marked exchange-rate turbulences.

These findings from aggregate data seem in contrast with accounts of national macroeconomic events of the last decade. The convergence that led to the European Monetary Union was a difficult and uncertain process. In at least the early part of it, some countries’ domestic real interest rates commanded substantial risk premia, while inflation expectations remained stubbornly high. Throughout the 1990s the process of macroeconomic convergence towards EMU took place in countries where initial monetary conditions and policy credibility were relatively similar (France) or relatively different (Italy) from those prevailing in Germany. During that period the countries belonging to the ERM went through a gradual hardening of the exchange-rate constraint. The "hard ERM" phase followed an initial crawling-peg-like experience, and the "wide band" arrangements that were enacted just after the 1992-93 crisis. The Taylor rule tries to capture the systematic component of monetary policy decisions. It is therefore somewhat puzzling that a euro-area-wide version of it turns capable to fully track interest-rate behaviour in such a changing context. We therefore estimate single-country instrument rules, and adopt a procedure that allows for a smooth transition model of policymakers’ reaction to the economy. In this way, we should be able to identify the effects of policy changes on single parameters of the policy rule. We adopt the class of simple reaction
functions we illustrated above, and measure the impact of policy change on interest-rate rules by studying how the coefficients of those functions vary over time. Our estimated interest-rate rules therefore have time-varying parameters. We consider five advanced countries: the US, the UK, France, Germany, Italy.

We employ a more flexible procedure than conventional approaches also because we believe that policymakers’ behaviour could display nonlinearities and asymmetries (Cukierman and Gerlach, 2003; Dolado et al., 2004). On the one hand, the central bank may react differently to inflation shocks in relatively high and low inflation regimes. On the other hand, it could be more sensitive to output below than above potential. Finally, Smets (2002) shows that the extent of uncertainty about economic conditions affects the coefficients of the Taylor rule. Since there is evidence of a worldwide decline in business cycle volatility over the 1980s and 1990s (McConnell and Perez-Quiros, 2000; Stock and Watson 2003), this might have had an effect on interest-rate rules that would be interesting to evaluate with some precision.

Next Section details our approach.

3 Gradual Policy Change: A TVP Approach

The models we estimate drop the constant coefficients hypothesis implicit in existing estimates of interest-rate rules. We allow our coefficient vector to change over time applying the Kalman filter algorithm to estimate the model. The estimation of models with time-varying parameters was pioneered by Doan, Litterman and Sims (1984). In this paper we follow their Bayesian approach, which allows the parameters of the models to evolve as more observations are added. The intuitive appeal of this approach is that policy regime and structural changes can be modeled as a gradual evolution of the model’s coefficients. The above requires to represent the model in the following general state-space form (see Harvey, 1989, and Kim and Nelson, 1999):

\[ i_t = c_t + x_t' b_t + e_t \]  

\[ b_{t+1} = d + T_t b_t + z_t \]  

where \( e_t \sim N(0, \sigma^2) \), \( z_t \sim N(0, Q) \), and \( b_0 \sim N(a_0, \Sigma_0) \)

[12] The description here closely follows the notation used in Hamilton (1994).
with $x_t$ containing the explanatory variables.

Equation (5) is the so-called *measurement or observation equation*. It is the classical linear regression model except that the parameter vector $b_t$ (representing the state variables) is allowed to stochastically change over time according to the *transition equation* (6). We follow Doan et al. (1984) in postulating a Bayesian prior distribution for the first-period value of the coefficient vector. Note that, in our model, the unobserved state vector $b_t$ is assumed to change over time according to a first-order vector autoregressive process. This prior distribution holds that changes in the endogenous variable modelled are so difficult to forecast that the coefficient on its lagged value is likely to be near unity, while all other coefficients are assumed to be near zero. The prior distribution is independent across coefficients, so that the MSE of the state vector is a diagonal matrix. Measurement errors and the disturbances to transition equations are assumed to be serially and mutually independent. The initial conditions (i.e the prior distribution) for the state vector $b_t$ are specified in (7); the elements of $Q$ are estimated through ML methods along with the other parameters of the model.

Summing up, this time-varying formulation involves forecasting the optimal state vector in each period, based on information available up to the previous period. Under the normality and independence assumptions about the disturbances, the computation of the state vector is simply obtained by applying the Kalman filter (see Hamilton, 1994). This way we obtain filtered estimates of the parameters and the residuals for each observation in the sample, thus accounting for the potential variation over time of the underlying structural parameters. This allows us to capture major regime shifts.

The estimated functions we show results for are those which yielded the best fit amongst the various specifications we tried. The sample period was 1980 to 1998 for Germany, France, Italy and the UK, whereas for the US we use data spanning 1970 to 2003. We stopped the sample to the establishment of the European single monetary policy, as this marked the end of national monetary policies for three of the countries in our sample. We inserted the UK in the group of countries we study not only for his macroeconomic links with the other countries, but also because its monetary policy witnessed operational evolutions partly shared by other economies. Last, but not least, the UK adhered to the ERM between 1990 and 1992.

Our starting model was in each case equation (4). This formulation is consistent (see Woodford, 2003) with forward-looking specifications of the
\[ i_t = \alpha + \beta E_t \pi_{t+j} + \gamma (y - y^*)_{t-s} + \sum_{n=1}^{k} \rho_n i_{t-n} + \sum_{p=1}^{l} \delta_p x_{t-p} + \varepsilon_t \]  

(8)

in which, for simplicity, we have not included time subscripts for the coefficients.

In line with common practice (see Clarida et al., 1998), we also estimated alternative specifications with additional variables \((x_t)\) such as measures of the exchange rate and a relevant foreign interest rate. This allows to check whether monetary authorities targeted intermediate or final objectives over and beyond those included in the baseline specification. As to the output term, Woodford (2003) acknowledges that the output variable that properly appears as an argument of an optimal policy rule is the difference between real GDP and a target level. This variable will change in response to real disturbances (preferences, technology, fiscal policy). However, the evolution of all these real factors could not follow the smooth trends (HP-derived) that literature commonly employs.

To measure inflation expectations and the output gap various methods are available. For inflation it is common to find measures of trend inflation extracted through the application of Hodrick-Prescott (HP) and band-pass filters; alternatively, results from market surveys are available. For the measurement of output gap, the production function approach and various multivariate methods provide more sophisticated alternatives to available filtering methods. The use of the HP filter to proxy the output gap is popular in applied work, although it is certainly not exempt from criticism. Estimates do not take into account the existence of potential measurement errors, nor do they take into account the two-sided nature of the filter, which may cause violations of weak and strong exogeneity assumptions. In the case of the US only, we computed the output gap series using the data on potential GDP provided by the Congressional Budget Office\(^{13}\).

For the remaining countries, we employ the Structural Time Series approach proposed by Harvey (1989) to generate series for the output gap. The procedure amounts to decomposing the series into trends, recursive stochastic cycles, and irregular (surprise) components that vary over time. This way, we fit univariate models for real GDP for each country, and extract time-varying measures of potential output that for each observation rely only on information available up to the point of estimation. This modelling

\(^{13}\)We compared this series with what one obtains from the application of the Kalman filter to actual GDP. The differences between the two are minor.
approach too applies a Kalman-filter estimation procedure. In the case of both series, the technique allows to assume a plausible learning process for both the central bank and private agents.

We measured expected inflation by extracting a trend from the original inflation series using the HP filter. This permits us to compare our results with those of most existing studies. Competing procedures, like using alternative filtering methods or polynomial trends, did not yield a significantly better fit of the estimated policy rules.

As to the coefficients’ lags, we report estimates for equations with significant coefficients only. Typically, we found that a lag length of \( n = 1 \) was sufficient to capture the inertial behaviour of the policy instrument induced by monetary authorities. As to the lead for the inflation regressor, we performed a search for it on the simple basis of goodness-of-fit criteria. In most cases the reaction lead was decided by statistical significance; in one or two cases by the ease of convergence of the optimization algorithm. For this reason, we found that when using filtered measures of expected inflation, the appropriate lead was between \( j = 2 \) and \( j = 4 \) quarters. In other instances, when the best fit was associated with actual future inflation, thus a more Taylor-like specification, the longest lead was 2 quarters\(^{14}\). These regularities are in line with most dynamic simulations of calibrated theoretical models (see Taylor, 1999; Woodford, 2003).

As is well known, the lack of statistical significance for the output term in a policy rule does not imply that interest-rate policy is not geared at stabilising output shocks. Indeed, monetary authorities can target output developments by simply adjusting the responsiveness of policy rates to inflation shocks and their overall degree of inertia. However, in most of our estimates one lag of the dependent variable, and the output gap, either lagged or current\(^{15}\), turned out to be significant.

4 Estimation Results

This Section presents our findings for the five countries in our sample. For each country we show estimates of the coefficients in the interest-rate reaction function (8). The presence of a \( t \) subscript indicates that we assume the coefficient as varying stochastically over time according to stationary

\(^{14}\)For brevity, here we do not report those estimates, which are anyway available on request.

\(^{15}\)In each case we show the specification for which the output gap turns out to be statistically significant.
first-order autoregressive processes with drift. Therefore, in terms of the instrument rules we estimated, the elements of vector $T$ in equation 5 obey to

$$
\begin{align*}
\beta_t &= c + \phi \beta_{t-1} + \eta_t \\
\gamma_t &= d + \lambda \gamma_{t-1} + \zeta_t \\
\rho_t &= h + \theta \rho_{t-1} + \mu_t \\
\delta_t &= f + \tau \delta_{t-1} + \omega_t
\end{align*}
$$

where and all errors terms are zero-mean, constant variance disturbances as detailed above. In some cases the specification with the best fit implied time-invariance for single coefficients; in these cases a time subscript does not appear.

The main features of our findings are that:

1. interest-rate rules diverge widely across our sample, e.g., they summarize policy conducts across countries according to different coefficient estimates, and, in some cases, different functional forms;

2. coefficient estimates seldom display statistically constant behaviour.

These findings vindicate the choice of estimating single-country interest-rate rules, and the validity of our TVP approach.

We now present our main results for each country.

4.1 USA

For the USA (sample 1970:1 to 2003:4) the specification with the best fit refers to the following interest-rate reaction function:

$$
i_t = \alpha + \beta_t E_t \pi_{t+4} + \gamma_t (y - y^*)_{t-1} + \rho_t i_{t-1} + \varepsilon_t
$$

where $i_t$ is the nominal Federal Funds Rate, $E_t \pi_{t+4}$ is 4-quarter ahead annualized expected inflation, computed by applying the standard HP filter to the inflation series, and $y - y^*$ is the output gap.

Figure 1, left-hand column, presents charts with smoothed estimates of the state vector\textsuperscript{16} $(\beta_t, \gamma_t, \rho_t)$, accompanied by confidence bands. Table 1

\textsuperscript{16}As is well known, it is possible to obtain filtered estimates of the state vector (Harvey, 1989). In all our cases the differences between the two sets of estimates were minor, so we report only smoothed values.
shows, for each state variable, the minimum, maximum, and mean values of the coefficients in the sample, and the date they occurred.

<table>
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<th>state</th>
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<th>max</th>
<th>mean</th>
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<td>$\beta_t$</td>
<td>2001:4</td>
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<td>0.19</td>
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<tr>
<td></td>
<td>(0.09)</td>
<td>(0.12)</td>
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<tr>
<td>$\gamma_t$</td>
<td>1984:4</td>
<td>2000:3</td>
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<td>0.09</td>
<td>0.35</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.11)</td>
<td>(0.10)</td>
<td></td>
</tr>
<tr>
<td>$\rho_t$</td>
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<td>1980:4</td>
<td>0.608</td>
</tr>
<tr>
<td></td>
<td>0.20</td>
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<tr>
<td></td>
<td>(0.04)</td>
<td>(0.06)</td>
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</table>

Table 1, USA, 1970:1-2003:4. State variables minimum, maximum, and mean values. Root mean squared errors in parenthesis.

Our results largely coincide with narrative accounts of Fed’s policy. The response of policy rates to inflation rises significantly since around 1977-78, peaking in 1984. Afterwards, the response to inflation declines, with a temporary correction around mid-1990s, after which it displays point estimates closer to the values it has for early 1970s. It is interesting to note that our estimates track quite well the generally loose monetary policies followed since 1995, which characterized Fed’s conduct around the so-called "productivity miracle". The behaviour of the coefficient on output gap is pretty symmetrical. The coefficient is statistically significant for most of the sample period, with the exception of the years overlapping with Volcker’s policy tightening. On the other hand, interest-rate smoothing seems to be a systematic feature of Fed’s policies: the relative coefficient is significant over the entire sample period.

The charts on the right-hand column of Figure 1 show the behaviour of the coefficients on inflation ($\beta_L$) and output gap ($\gamma_L$) derived from the long-run solution of our estimated interest-rate reaction function. The long-run static reaction function, obtained by inserting the mean values of $\alpha_L = \hat{\alpha}/(1 - \hat{\rho}_t)$, $\beta_L = \hat{\beta}_t/(1 - \hat{\rho}_t)$, and $\gamma_L = \hat{\gamma}_t/(1 - \hat{\rho}_t)$ into the short-run estimated equation, is

$$i_t = 1.55 + 1.28E_t\pi_{t+4} + 0.69(y - y^*)_{t-1} + \varepsilon_t$$
The static long-run policy rule\textsuperscript{17}, based on our time-varying estimates, points to a stabilizing reaction of policy rates to inflation (the inflation coefficient exceeds unity), and a distinctive break in 1980. Coefficients’ point estimates reflect the volatility of short-term interest rates subsequent to the temporary adoption of money growth targets based on bank reserves in October 1979. Therefore, our estimates appear broadly in line with standard Taylor-rule values obtained elsewhere. This confirms that, while capable of portraying a richer picture of the evolution of policy, our techniques are consistent with theoretical approaches to policy evaluation.

4.2 Germany

For Germany (sample 1980:1 to 1998:4) we show results for the following reaction function

\[ i_t = \alpha_1 + \alpha_2 q_t + \beta_t E_t \pi_{t+4} + \gamma_t (y - y^*_t) + \rho_t i_{t-1} + \delta_t i_t^{\uparrow} + \varepsilon_t \]

in which we inserted, along with regressors that we have already introduced, a measure of the real effective (trade-weighted) exchange rate \( (q_t) \), while \( i_t^{\uparrow} \) is the US Federal Funds rate\textsuperscript{18}.

The graphs in Figure 2 show short-run coefficients that move relatively little. This suggests that the underlying policy objectives were more or less stable. Table 2 shows that the short-run coefficient on the output gap is barely significant. However, its estimate should be evaluated along the statistically significant and sizeable inertia in policy rates. Overall, this confirms that Bundesbank’s policies systematically followed cyclical conditions, despite being formulated in terms of explicit monetary targets. Finally, the real exchange rate enters significantly (p-value: 0.0004), pointing out that external conditions too were relevant in shaping the behaviour of domestic policy rates. However, the coefficient on the Federal funds rate is not significant: apparently, European macroeconomic developments held the upper hand in determining the policy stance.

\textsuperscript{17}It would be helpful to have an idea of the statistical significance of the long-run static coefficients. However, we are not aware of simple procedures to obtain such “long-run standard errors” in the context of our approach. We could obviously resort to Monte Carlo methods or similar techniques, but they would surely make our simple exercise much more cumbersome.

\textsuperscript{18}Note that the estimated specification features a constant coefficient on the real exchange rate and a time-varying one on the Federal funds rate. We experimented with alternative specifications, but the one we present was the best-fitting one.

<table>
<thead>
<tr>
<th>state</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta_t$</td>
<td>1994:2</td>
<td>1992:1</td>
<td>0.50</td>
</tr>
<tr>
<td></td>
<td>0.42</td>
<td>0.60</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.05)</td>
<td>(0.05)</td>
<td></td>
</tr>
<tr>
<td>$\gamma_t$</td>
<td>1991:2</td>
<td>1982:4</td>
<td>0.16</td>
</tr>
<tr>
<td></td>
<td>0.06</td>
<td>0.26</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.05)</td>
<td>(0.05)</td>
<td></td>
</tr>
<tr>
<td>$\rho_t$</td>
<td>1994:2</td>
<td>1992:1</td>
<td>0.73</td>
</tr>
<tr>
<td></td>
<td>0.71</td>
<td>0.76</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.02)</td>
<td></td>
</tr>
<tr>
<td>$\delta_t$</td>
<td>1985:3</td>
<td>1997:4</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>-0.05</td>
<td>0.05</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.02)</td>
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</tbody>
</table>

The long-run static solution to the reaction function,

$$i_t = -0.19 - 23.33q_t + 1.89E_t \pi_{t+4} + 0.59(y - y^*)_t + 0.002i_t^{IF} + \varepsilon_t$$

yields point estimates that are consistent with a strong antinflationary commitment (the inflation coefficient is well above unity). Figure 2 shows that the Bundesbank’s policy thrust turned even more conservative over time, and specifically after 1991, when the central bank tried to offset the inflationary consequences of the reunification with Eastern Landers.

4.3 UK

For the UK (sample 1980:1 to 1998:4) we present estimates for the following interest-rate reaction function:

$$i_t = \alpha + \beta_t E_t \pi_{t+4} + \gamma_t (y - y^*)_t + \rho_t i_{t-1} + \delta_t i_t^{IF} + \varepsilon_t$$

where, alongside already-introduced regressors, we inserted the 3-month FIBOR German rate (i_t^{IF}). We did this to take into account the evolving importance that external constraints played in the conduct of monetary policy.

19We do not present the estimates we obtained for models featuring the real exchange rate (alternatively measured), as the relative coefficients were never significant.
policies in the UK. In particular, the linkage with Bundesbank’s policies emerges very clearly from existing narrative and econometric analyses (see for instance Nelson, 2002; Muscatelli et al., 2002b).

Results in Table 3 and Figure 3 show indeed that evaluating UK’s monetary policy only on the basis of the estimate of the coefficient on inflation would be quite restrictive. That coefficient turns out to be significant, but so too are the coefficients attached to interest-rate smoothing and FIBOR. In particular, they both display an interesting short-run dynamics, as they capture sharp tightenings around mid-eighties. That period saw the abandonment of the loose monetary targeting regime implemented since the late 1970s, and a renewed emphasis on exchange rate targeting (the "shadowing the D-Mark" phase). The coefficient on output gap is rarely significant, but, given the significance of the interest-rate smoothing term, this does not rule out a systematic emphasis of the Bank of England on output stabilisation.

<table>
<thead>
<tr>
<th>βt</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>1985:2</td>
<td>0.42 (0.06)</td>
<td>0.77 (0.06)</td>
<td>0.62</td>
</tr>
<tr>
<td>1985:3</td>
<td>0.77 (0.06)</td>
<td>0.77 (0.06)</td>
<td>0.62</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>γt</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>1988:3</td>
<td>-0.15 (0.13)</td>
<td>0.58 (0.23)</td>
<td>0.30</td>
</tr>
<tr>
<td>1981:3</td>
<td>0.58 (0.23)</td>
<td>0.58 (0.23)</td>
<td>0.30</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>δt</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>1994:4</td>
<td>0.27 (0.05)</td>
<td>0.81 (0.08)</td>
<td>0.49</td>
</tr>
<tr>
<td>1985:2</td>
<td>0.81 (0.08)</td>
<td>0.81 (0.08)</td>
<td>0.49</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>ρt</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>1985:2</td>
<td>0.34 (0.04)</td>
<td>0.35 (0.04)</td>
<td>0.34</td>
</tr>
<tr>
<td>1986:4</td>
<td>0.35 (0.04)</td>
<td>0.35 (0.04)</td>
<td>0.34</td>
</tr>
</tbody>
</table>


Long-run static estimates confirm this behaviour,

\[ i_t = 1.16 + 0.95E_t \pi_{t+4} + 0.45(y - y^*)_t + 0.75\delta_{t,1} + \varepsilon_t \]

where the most important feature to notice is not the solved value of the βt coefficient (slightly below unity), but the magnitude of the coefficient associated with Germany’s FIBOR.
4.4 France

For France (sample 1980:1 to 1998:4) the specification with the best fit was

\[ i_t = \alpha + \beta_t E_t \pi_{t+4} + \gamma_t (y - y^*)_t + \rho_t i_{t-1} + \delta_t i^f_t + \varepsilon_t \]

We also experimented with a rule that featured a measure of the real effective (trade-weighted) exchange rate, but we found it never significant. Since it was also the cause of non-convergent estimation, we dropped it altogether.

The charts in Figure 4 and Table 4 show that the early 1980s saw a sharp increase in the response of policy rates to changes in inflation. Such reaction appears to have remained more or less constant throughout the rest of the sample. The optimal lead on expected inflation turns out to be the same as in Germany \((j = 4)\).

<table>
<thead>
<tr>
<th>state</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\beta_t)</td>
<td>1980:3: 0.18 (0.05)</td>
<td>1984:1: 0.43 (0.06)</td>
<td>0.34</td>
</tr>
<tr>
<td>(\gamma_t)</td>
<td>1984:1: 0.04 (0.14)</td>
<td>1997:4: 0.53 (0.09)</td>
<td>0.22</td>
</tr>
<tr>
<td>(\rho_t)</td>
<td>1993:2: 0.40 (0.02)</td>
<td>1993:1: 0.48 (0.02)</td>
<td>0.44</td>
</tr>
<tr>
<td>(\delta_t)</td>
<td>1980:4: 0.21 (0.06)</td>
<td>1995:2: 0.63 (0.05)</td>
<td>0.33</td>
</tr>
</tbody>
</table>


Against this background, French interest rates too displayed significant inertia, and a growing sensitivity to output-gap developments, as shown by charts in the right-hand column of Figure 4. Finally, the long-run static solution to the reaction function,

\[ i_t = 3.37 + 0.61 E_t \pi_{t+4} + 0.39 (y - y^*)_t + 0.59 i^f_t + \varepsilon_t \]

\(20 i^f_t\) is the 3-month FIBOR rate.
shows that, contrary to previous cases, the long-run response of rates to inflation falls short of one. Nonetheless, the estimates confirm that French rates (as in the case of the UK) displayed a systematic tendency to follow German interest rates, whose presence in the function clearly subtracts significance to the inflation coefficient.

4.5 Italy

The estimation of a reaction function for Italy posed the most severe convergence problems. Various configurations of the policy rules were evaluated, but convergence of our optimization algorithm was hard to obtain, and estimates seldom provided significant coefficients. Moreover, results often proved too sensitive to small changes in the sample period or lag structure. This is again compelling evidence that a "one-size-fits-all" approach to this investigation would not necessarily be able to explain relevant features of each country’s monetary policy. In the case of Italy, specifications that included inflation expected four quarters in advance did not emerge as so fitting as for the other countries. Since the model with the best fit was one in which the lead on expected inflation was only two quarters, we argue that monetary policy in Italy tended to be less forward-looking than in the other countries. We show results for the following reaction function over the usual sample 1980:1 to 1998:4:

\[ i_t = \alpha_1 + \alpha_2 i_{t-1} + \beta_t E_t \pi_{t+2} + \gamma_t (y - y^*)_t + \delta_t i_t + \epsilon_t \]

where \( E_t \pi_{t+2} \) is 2-quarter-ahead annualized expected inflation, computed by applying the standard HP filter to the inflation series.

The clearest result emerging from Table 5 and Figure 5 is that the Bank of Italy targeted both domestic inflation and German interest rates. The interest-rate-smoothing coefficient has a point estimate of 0.43, but it is not significant at the conventional 95% confidence level (s.e. = 0.23). The output gap enters significantly only in the latter part of the sample, when, on the other hand, the degree of aggression towards inflation somewhat
increases.

<table>
<thead>
<tr>
<th>state</th>
<th>min</th>
<th>max</th>
<th>mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta_t$</td>
<td>1981:1, 0.45 (0.06)</td>
<td>1986:1, 0.69 (0.06)</td>
<td>0.54</td>
</tr>
<tr>
<td>$\gamma_t$</td>
<td>1980:2, -0.32 (0.22)</td>
<td>1998:2, 0.55 (0.17)</td>
<td>0.10</td>
</tr>
<tr>
<td>$\delta_t$</td>
<td>1998:2, 0.06 (0.08)</td>
<td>1992:3, 0.60 (0.04)</td>
<td>0.22</td>
</tr>
</tbody>
</table>


This dynamics is confirmed by the coefficient values of the long-run static solution to the reaction function,

$$i_t = 4.35 + 0.95E_t\pi_{t+2} + 0.18(y - y^*)_t + 0.39i_{t-1}^f + \varepsilon_t$$

All in all, the evolution of coefficients is consistent with the historical analysis regarding Italian monetary policy\(^{21}\). Specifically, for the years 1981-86, the rise in the $\beta$ estimate accords with the view that, after the "divorce"\(^{22}\) Italian monetary policy shifted to a more inflation-averse stance. The Bank of Italy gained further independence during 1992-93\(^{23}\). With increasing independence, the likely shift in the Bank’s toward a more inflation-averse stance is behind the increase of $\beta$ observed around those years. Finally, commitment to ERM forced Italian monetary policy to pay growing attention to the external constraint. In fact, estimates of $\delta$ are significant and

\(^{21}\)See Fratianni and Spinelli (2001).

\(^{22}\)So was defined the relationship between the Italian government and the Bank of Italy following the law (June 1981) abolishing the obligation to buy unsold Treasury Bonds by Bank of Italy.

\(^{23}\)The most important innovations were the fixing of the official discount rate by Bank of Italy alone, without any ratification by the Treasury (30/01/1992) and the abolition, dating 12/11/1993, of the use of the Treasury’s account at BI to finance public deficit (it was possible up to the 14 % of its amount). Also the new rules (9/02/1993) regarding the obligatory reserves of the banking system at BI were of great importance. In fact, even after divorce the BI would use that liquidity to buy Treasury bonds at non-market conditions.
with the expected sign over the all sample. The graphs show that the external constraint was particularly binding during the two speculative attacks to the lira exchange rate at the end of 1985 and during 1992.

4.6 TVP Rule vs. Taylor Rule

In this subsection we briefly look at how our TVP-based rule performs vis-à-vis the best-known proposal for the tracking of actual policy rates, i.e., the Taylor rule. In particular, we feed the values of expected inflation, output gap and lagged policy instrument into our TVP reaction function, so that we obtain a series for the policy rate "implied" by our rule. We perform analogous computation using the rule proposed by Taylor (1993), which has swiftly become the benchmark for policy analysis and a consensus tool for forecasts by market participants:

\[ i_t = 4.00\% + 1.5(\pi_t - 2.00\%) + 0.5(y - y^*) \]

where the constants conventionally account for an implicit inflation target of 2 percent per annum, and an estimate of the long-run real policy rate of 2 percent (for the US) as well.

Figure 6 gives a visual idea of how closely the two rules track the behaviour of historical policy rates in the US over the period 1983-2003. Table 6 shows three simple measures of the relative accuracy of TVP and CT rules in following the developments of actual rates.

The TVP rule clearly emerges as superior for this purpose: the conventional Taylor rule (CT) underestimates the actual rate for most of the 1980s and again in the second half of the 1990s, whilst it overshoots the Fed funds rate in '92-'94 and 2000-2002. The TVP rule follows much more closely the actual behaviour of the short-term rate: its correlation with the Fed funds rate is 0.93, whilst the CT rule and the actual rate have a correlation of 0.55. Our rule’s main shortcoming is its inability to exactly pick up some loosening episodes of policy rates, as in 1984, 1989 and, to a lesser extent, 1995. On the contrary, tightenings appear as precisely tracked. This asymmetry possibly points to interesting nonlinearities in the response of monetary authorities to business cycle developments.

\[ ^{24}\text{Clarida et al. (1998) conclude that monetary policy in Italy (but the same was true for France and the UK) boiled down to fighting inflation by following the Bundesbank.} \]

\[ ^{25}\text{Obviously, this implies that keeping the long-run average inflation rate at the target requires a long-run average policy rate of 4 percent.} \]

\[ ^{26}\text{We do not report here results for the remaining countries for sake of brevity. However, our findings for those cases broadly reflect what we show for the US, and are available from the authors upon request.} \]
We also compute a) the mean value of deviations (ME), b) the mean value of absolute deviations (MAE), and the root mean squared error (RMSE) of each rule’s implied Fed funds rate from the actual data.

<table>
<thead>
<tr>
<th>Rules</th>
<th>ME</th>
<th>MAE</th>
<th>RMSE</th>
</tr>
</thead>
<tbody>
<tr>
<td>TVP</td>
<td>0.0524</td>
<td>0.7171</td>
<td>1.0097</td>
</tr>
<tr>
<td>CT</td>
<td>-0.7100</td>
<td>1.4488</td>
<td>2.0853</td>
</tr>
</tbody>
</table>

Table 6, U.S., 1983:1-2003:4. Mean value of deviations (ME), mean value of absolute deviations (MAE), and root mean squared error (RMSE) of the Fed funds rate implied by TVP and CT rules implied from the actual data.

The computed values confirm that our TVP rule outperforms the standard Taylor rule. In particular, the large negative value of ME states that the CT rule tends to underestimate the actual rate. MAE and RMSE values for CT are more than twice those of TVP. Obviously, the absence in CT of a smoothing term could partly account for this finding. However, our proposal appears richer than the CT on a variety of grounds. First, it allows for forward-looking behaviour of the policy maker in setting interest rates. Second, time-varying coefficients for expected inflation, output gap and the lagged interest rate provide a fuller picture of monetary policy shifts than any constant-coefficients approach could support. The latter feature in particular does not impair the descriptive power of our rule, given that it also performs well in tracking actual policy rates.

5 Conclusions

In this work we estimated forward-looking interest-rate rules for five major economies, using a Bayesian approach that allows for time variation in estimated parameters. Traditional constant-parameter instrument rules likely blur the impact of a series of relevant features of modern policymaking, like model uncertainty, conflicting objectives, shifting preferences and nonlinearities in policymakers’ behaviour. Our results instead frame monetary policy conducts in a dynamic and evolving context. The aim was to capture monetary policy shifts otherwise missed by constant-parameters approaches to interest-rate policies. This helps us to describe monetary policies in a richer way than conventional analyses, and to identify at least three general issues.

First, estimated policy rules summarize policy conducts according to different coefficient estimates, and, in some cases, different functional forms.
This implies that interest-rate policies diverge widely across countries. Second, estimated parameters tend to shift over time, in most cases in a smooth and gradual fashion. We believe that structural approaches, like DSGE and time-varying VAR models, will be able to assess the extent to which changes in preferences or shifts in "deep" parameters account for this evolution.

Finally, given the widespread interest in the evaluation of ECB policy conduct so far, our work sheds some light on the degree to which any average measure of policy rates in the euro area pre-1999 should be used in such exercises. As the support of our findings for a common Taylor rule for Germany, France and Italy appears very weak, some caution is in order.
Data Appendix

The data we used were quarterly series, extracted from OECD Main Economic Indicators, Thomson Financial Datastream, and IMF’s International Financial Statistics. In all cases we were able to employ seasonally adjusted data.

For each country, we measured output using GDP at constant price series. The output gap is defined as quarter-on-quarter log-difference between actual and potential levels of the series. The inflation series were defined as 4-quarter log-differences in the all-items CPI. The index of effective exchange rates (trade weighted) was the measure for the exchange rates.

The following is a short description of the variables’ sources.

- United States. The output series is the Real Gross Domestic Product, in Billions of Chained 2000 Dollars (source: U.S. Congress, Congressional Budget Office). The output gap is obtained using the Real Potential Gross Domestic Product, defined in Billions of Chained 2000 Dollars, same source as actual output. Inflation is the 4-quarter (log) difference in the Gross Domestic Product Chain-type Price Index, 1996=100, Seasonally Adjusted (source: U.S. Department of Commerce, Bureau of Economic Analysis). The call money rate is the Federal Funds’ rate, obtained from IMF’s IFS.

- Germany, France, United Kingdom and Italy. IMF’s International Financial Statistics (Call Money Rate, Consumer Price Index and Gross Domestic Product).

The short-term interest rate employed as the monetary policy indicator was the following:

- Germany, overnight money market rate
- United Kingdom, Overnight Interbank Rate series post-1983; pre-1983 we employ the Rate on 90-day Treasury Bills, which displays a very close correlation with the interbank lending rate (source: IMF, IFS)
- France, overnight money market rate
- Italy, 3-Month Interbank Deposits (Overnight)
References


Monetary Economics, 46: 281-313.


Fratianni, M., and F. Spinelli, 2001, Storia Monetaria d'Italia, ETAS.


Graph Appendix

Figure 1, USA. State variables smoothed estimates (left-hand column) and long-run static coefficients (right-hand column).
**Figure 2, Germany.** State variables smoothed estimates (left-hand column) and long-run static coefficients (right-hand column).
Figure 3, UK. State variables smoothed estimates (left-hand column) and long-run static coefficients (right-hand column).
Figure 4, France. State variables smoothed estimates (left-hand column) and long-run static coefficients (right-hand column).
Figure 5, Italy. State variables smoothed estimates (left-hand column) and long-run static coefficients (right-hand column).
Figure 6, US. Actual Fed funds rate (FFR), and implied Fed funds rate from time-varying parameter (TVP) and Taylor rules (CTR), 1983-2003. (see main text for details).
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