Bonn Econ Discussion Papers

Discussion Paper 21/2003

Monetary Policy in Europe: Evidence from Time-Varying Taylor Rules

by

Katrin Wesche

October 2003

Bonn Graduate School of Economics
Department of Economics
University of Bonn
Adenauerallee 24 - 42
D-53113 Bonn

The Bonn Graduate School of Economics is
sponsored by the

Deutsche Post World Net

MAIL EXPRESS LOGISTICS FINANCE

Monetary Policy in Europe
Evidence from Time-Varying Taylor Rules

Katrin Wesche
Institut für Internationale Wirtschaftspolitik
Universität Bonn
Lennéstraße 37, D-53113 Bonn

October 2003

Abstract

We estimate monetary policy reaction functions for France, Germany, Italy, the United Kingdom, and the United States using a Markov-switching model that incorporates switching in the monetary policy regime as well as an independent switching process for shifts in the state of the economy. Results indicate that over time all central banks have assigned changing weights to inflation and the output gap. Regimes can be classified as "dovish" with a high weight on output and a low weight on inflation, and "hawkish" with a high weight on inflation and a low one on output. For France and Italy, the German interest rate had an influence on domestic monetary policy especially at the beginning of the 1980s after the inception of the European Monetary System (EMS). Switching in the residual variance of the monetary rule accounts for heteroscedasticity and turns out to be important for the fit of the model. Robustness of the results is checked by considering alternative specifications of expected inflation and the output gap. In general, results are robust to these changes.

Keywords: Monetary policy rule, Taylor rule, Markov switching
JEL classifications: E41, E58, C22

The analysis of so-called monetary policy rules has become a widely used tool to assess a central bank's monetary policy. Though most central banks reject to obey monetary policy rules as an orientation for day-to-day monetary policy,[1] rules permit an ex-post evaluation of monetary policy in terms of a few, economically relevant variables. In this regard, the term "rule" used in the literature is misleading because the empirical analysis of monetary policy rules is concerned with a descriptive analysis of central bank behavior. Speaking of a monetary policy equation or a central bank reaction function thus would be more appropriate, but in accordance with most of the literature also here the term "rule" is used. By estimating a reaction function for the central bank, monetary policy rules allow to analyze the implicit goals of the central bank by determining the weight the central bank assigns to different economic indicators. In this respect, estimated monetary policy reaction functions can prove especially useful for the comparison of different policies across countries or over time.

One of the most widespread rules in the literature is the Taylor rule (Taylor 1993), which assumes that the central bank reacts to deviations of inflation and output from target. Taylor proposed his rule as a description of how the Federal Reserve Bank (FED) conducts monetary policy and showed that his rule closely tracks the actual federal funds rate in the United States from 1987 to 1992. Estimation of monetary policy reaction functions for other periods and other countries yielded similar results.[2] Moreover, Taylor-type rules have shown to be a useful formulation of monetary policy in simulations of different macroeconomic models.[3]

Empirical studies of monetary policy rules are typically confronted with the problem that central-bank policy changes over time, especially when considering a longer sample period.[4] According to conventional wisdom, central banks in the 1970s put a high weight on output while inflation was allowed to rise, whereas in the more recent past most central banks concentrated on achieving low inflation, and output goals received less attention. The central bank also may respond differently to economic variables depending on the values they take in a particular situation, i.e., the central bank may react asymmetrically to high and low values of inflation or output. Additionally, external constraints have changed over time and may have enforced

[1] One main problem with monetary policy rules like the Taylor rule for actual policy formulation is the imprecision of filtering methods used for output gap estimation at the end of the sample period, see e.g. Kozicki (1999), Deutsche Bundesbank (1999).

[2] See e.g. Clarida and Gertler (1997) for the Bundesbank, Clarida, Galí, and Gertler (1998) for the G3 and three large EU countries, Wyplosz (1999) for the EU countries, and Mihov (2001) for Germany, France, and Italy.

[3] See Levin, Wieland, and Williams (1999) or Christiano and Gust (1999).

[4] These changes may either take the form of a gradual shift or a sudden switch to another regime. While a gradual shift in general would be better modelled by smooth transition models, this paper is concerned with more or less abrupt changes of regime.

different policy reactions. While after the breakdown of the Bretton-Woods System exchange rates in the 1970s in most economies followed a free float, the 1980s saw the successive hardening of the exchange rate constraint for the member countries of the European Monetary System (EMS), which finally lead to the transition to European Monetary Union (EMU).

This paper is structured as follows. In Section 1 the existing literature on time-varying monetary policy rules is reviewed. Theoretical models based on asymmetric central-bank preferences imply a time-varying reaction to economic variables. Though empirical research on this topic is still in the beginning, non-linearities in monetary policy rules seem to be of importance. Section 2 discusses the specification of the Markov-switching model and the estimation procedure. It is assumed that the monetary policy regime switches according to a first-order Markov process, while a second independent Markov process determines switching in the residual variance of the monetary policy rule. The countries investigated are France, Germany, Italy, the United Kingdom and the United States. Section 3 discusses the data and Section 4 presents the estimation results. Results show that the weights assigned to inflation and the output gap follow two distinct regimes. The first monetary policy regime is associated with a high weight on inflation, while in the second regime the central bank follows an accommodative policy. Following the terminology of Owyang and Ramey (2000) one regime can be classified as "hawkish" because a high weight on inflation is associated with a low one on output, whereas the other regime can be termed as "dovish" with a high weight on output and a low weight on inflation. In addition, the consequences of participation in the EMS for France and Italy are explored. Switching in the residual variance is relevant for all countries and contributes significantly to an improvement over a simple linear model. Results are found to be robust against changes in the definition of expected inflation and the output gap. Section 5 concludes.

1 Time-Varying Monetary Policy Rules

In the literature various reasons have been discussed why central-bank reactions to economic variables may vary over time. By studying the minutes of the FOMC meetings, Romer and Romer (1989) find that at certain times the FED has been concerned about inflation and has been willing to take output losses. Blinder (1998, p. 19) states that political pressure on the central bank is higher when it tightens monetary policy preemptively than when it eases preemptively. Fischer (1994, p. 293) presumes that dependent central bankers suffer from an inflationary bias, while independent central bankers develop a deflationary bias. With increasing independence, the central bank's preference towards inflation presumably will shift. While the change in independence is less relevant for the U.S. and Germany, other countries, like the United Kingdom, France and Italy, have seen a fundamental shift in their central bank constitutions, which could well have influenced the monetary policy rule.

While inspection of central-banking practice confirms the possibility of time-varying responses to economic variables, also theoretical models give explanation for asymmetric reactions to target variables. Amano, Coletti, and Macklem (1999) show that central banks would have to adjust parameters of inflation forecast based rules to achieve low inflation and low output variability if the economic environment—e.g., central-bank credibility or the anti-cyclical properties of fiscal policy—changes. Additionally, uncertainty about the state of the economy may affect the monetary rule. Rudebusch (2001) and Smets (2002) argue that coefficients of a Taylor rule are affected by the imprecision of output gap estimates. As uncertainty about output growth and inflation in the 1970s after the oil-price shocks was presumably higher than in the 1990s, this could have altered the parameters of the monetary rule.

Moreover, central banks could react differently to positive and negative deviations of inflation or output from target, e.g., the central bank may react more aggressively to inflation if the economy is in a boom than if it is in a recession. Cukierman's (1999) model implies that the central bank's preference for inflation increases with the likelihood of a recession. Jordan (2001) shows that a monetary control error in combination with asymmetric preferences for output above and below target leads to higher inflation. Ruge-Murcia (2001) develops a model where inflation preferences of the central banker are asymmetric. If inflation is above target, the central bank reacts with a restrictive monetary policy, while an inflation rate below target is not countered by an equally expansive policy.

Also empirically time-varying behavior of central banks is supported. Exploiting the resulting non-linear relationship between inflation and unemployment from his model, Ruge-Murcia (2001) finds that for the United Kingdom, Sweden, and Canada the central bank seems to weight positive deviations of inflation from target more heavily than negative ones. Gerlach (2000) assumes a non-linear function for the output coefficient in a Taylor rule and finds an asymmetric output reaction for the FED during the period 1960–79, but no asymmetry for the second half of the sample from 1980 to 1999. Asymmetric reactions to inflation are also found by Dolado, María-Dolores, and Naveira (2000) for France, Germany, Spain, and the United States, using dummy variables for inflation above and below target. Finally, for the FED Clarida, Gali, and Gertler (2000), and Judd and Rudebusch (1998) split the sample period and show that during the last thirty years coefficients of the monetary policy reaction function have changed with different chairmen.

Nevertheless, the empirical studies discussed so far have some limitations. Splitting the sample period leads to a shortening of the available time series. Instead of adjusting the sample period, a switching model uses all available data and at the same time permits more flexibility than a linear model as coefficients can take different values in each regime. Moreover, adoption of a switching model has the advantage that one does not have to decide prior to the estimation which factors determine changes in central-bank policy. Thus one can "let the data speak" and then try to interpret the results in terms of estimated coefficients and the timing of regime switches. In this paper the base-line specification of the Markov-switching model characterizes central-bank policy as falling into two different regimes. One regime is expected to correspond to an anti-inflationary regime with a high weight on inflation and low one on output, while in the second regime the central bank should take a more accommodative position towards inflation with a high coefficient on output and a low one on inflation.

Switching models are also able to deal with the changing variability of economic time series. The oil-price shocks in the 1970s increased volatility of interest rates, inflation and output. Also the episode of base-money targeting in the U.S. from 1979 to the early 1980s induced considerable volatility in interest rates. Finally, the turbulences in the European Monetary System (EMS), especially the exchange rate crises of 1992/93, influenced monetary policy and the variance of the interest rate. As these episodes are mostly unrelated to the monetary policy regime in place, they are assumed to switch independently from switching process driving the coefficients in the monetary policy rule.

Up to now, few studies have investigated monetary policy reaction functions with time-varying coefficients. Sims (1999) estimates a 3-state switching model on a short-term interest rate and the consumer price index with simultaneous switching in the coefficients and the variance. He finds that different regimes in the disturbances help improve the fit of the model while different coefficients are less important. Owyang and Ramey (2001) estimate a model where the inflation target as well as the NAIRU switch independently between a low and a high state. They find that monetary policy in the United States has switched between an accommodative "dove" regime and a less accommodative "hawk" regime. To our knowledge, monetary policy rules with switching coefficients have not been estimated for European countries yet, except for Dueker and Fischer (1996) for Switzerland. It therefore seems interesting to investigate changing policy reactions for European countries in a Markov switching framework and to compare the outcome with the results for the United States.

2 The Model

In an influential article, Taylor (1993) proposed the following rule as a characterization of monetary policy.

$$i_t^T = \bar{r} + \pi_t^* + \beta(\pi_t - \pi_t^*) + \gamma(y_t - y_t^*) \tag{1}$$

Corresponding to the practice of most central banks the interest rate is regarded as the monetary policy instrument. Taylor assumes that the federal funds rate should rise if inflation, π, rises above target or if output, y, rises above its trend value. In equilibrium the deviation of inflation and output from their target values is zero and the desired interest rate, i^T, is the sum of the equilibrium real rate, $\bar{r}$, plus the target value of inflation, π^*. The last two terms of equation (1) show the influence of the deviation of inflation from its target value and the influence of the deviation of actual output from its trend value, y^*. Taylor assumed a value of 2 % for $\bar{r}$ and for π^*, a trend output growth of 2.2 %, and weights of 1.5 for β and 0.5 for γ. He showed that the actual interest rate policy of the FED from 1987 to 1992 has been conducted as if the FED had followed such a policy rule.

Instead of assuming specific values for the coefficients, the literature generally has estimated weights for the reaction function. To keep things simple, Taylor's (1993) original rule considers the deviation of inflation over the last four quarters from target. Most central banks, however, do not target past or actual but expected inflation. In the literature, therefore, in general forward looking—i.e., inflation forecast based—rules are considered.[5] This allows the central bank to take various relevant variables into account when forming its inflation forecasts.

Starting point for the empirical model is the following central bank reaction function developed by Clarida, Galí and Gertler (1998). The central bank's desired target interest rate i^T depends on the deviation of expected inflation j periods ahead from its target value, π^*, and the expected output gap k periods ahead; i^* is the equilibrium nominal interest rate.

$$i^T = i_t^* + \beta\left[E(\pi_{t+j}) - \pi_{t+j}^*\right] + \gamma\left[E(y_{t+k}) - y_{t+k}^*\right] \tag{2}$$

For monetary policy to be stabilizing, the weight on the inflation gap should exceed unity and the coefficient on the output gap should be positive. A coefficient greater than unity on the inflation gap means that the central bank pushes up the real rate in response to higher inflation, which exerts a stabilizing effect on inflation. A positive coefficient on the output gap entails lower interest rates in situations where output

[5]See Amano, Coletti, and Macklem (1999), Batini and Haldane (1999) or Clarida, Galí and Gertler (1998, 2000).

is below normal and thus has a stabilizing effect on the economy. By investigating monetary policy rules one thus can check if monetary policy has behaved adequately over time.

To account for the observed autocorrelation in interest rates, a dynamic adjustment for the interest rate is used.

$$i_t = (1-\rho)i_t^T + \rho \sum_{i=1}^{n} w_i i_{t-i} + \varepsilon_t \quad \text{with} \quad \rho \in [0,1], \quad \sum_{i=1}^{n} w_i = 1 \tag{3}$$

Equation (3) states that the central bank does not adjust the interest rate immediately to its desired level but is concerned about interest rate smoothing. Inserting equation (3) into equation (2) yields the following equation

$$i_t = (1-\rho)\left[\alpha + \beta E(\pi_{t+j}) + \gamma\left(E(y_{t+k}) - y^*_{t+k}\right)\right] + \rho \sum_{i=1}^{n} w_i i_{t-i} + u_t.$$

2.1 Non-linear Taylor Rules

In empirical applications, different forms of dealing with time variation in the parameters of a monetary policy reaction function have been used. One possibility is to split the sample at the presumed break date and estimate the equation for both periods separately. This approach has been taken by Clarida, Galí and Gertler (2000) or Judd and Rudebusch (1998), who investigate the effect of different central bank presidents on monetary policy for the FED, and also by Neumann and von Hagen (2002) who test for changes in the Taylor rule due to the introduction of inflation targeting in a sample of six countries.

Another approach is to use dummy variables to estimate different coefficients depending on inflation or output being above or below target (Dolado, María-Dolores, and Naveira 2000). The disadvantage of these methods is that one has to find exogenous information indicating that a switch in regime has occurred. Splitting the sample shortens the available time series. This is not attractive as in the estimation it is generally assumed that long-run inflation and the long-run real interest rate equal their equilibrium values, which is only the case if the sample period is sufficiently long. While this problem is avoided in the dummy variable approach, also here the researcher has to maintain a hypothesis what factors are responsible for a shift in regime.

The assumption of a non-linear function for the coefficients in a Taylor rule, as in Gerlach (2000), or for the relation between inflation and unemployment as in Ruge-Murcia (2001), is an interesting alternative to the approach pursued in this paper. In contrast to a sudden switch in regime, a non-linear function implies a gradual change in central-bank behavior as the respective variable deviates from target.

Besides the different nature of regime shifts modeled by a non-linear function and the assumption of Markov switching, in switching models no a-priori assumptions on the causes of the regime shifts or the functional form for the changes in regime have to be made.

Markov switching models first have been used in business-cycle and exchange-rate analysis (Hamilton 1989, Engel and Hamilton 1990). However, they also have a natural interpretation for monetary policy, as argued above, since monetary policy rules are unlikely to be constant over time. Markov switching is an attractive way to model the succession of different regimes where switching between regimes does not occur deterministically but with a certain probability. These regimes are not classified ex-ante but are estimated from the data. Apart from estimating how much weight the central bank puts on relevant economic variables like inflation and output, one can investigate how these weights change over time.

Assuming that the central bank's reaction to the deviation of inflation from target and the output gap depends on the monetary policy regime in place, the model equation becomes

$$\begin{aligned} i_t &= (1-\rho)\left[\alpha + \beta(S_{1,t})E(\pi_{t+j}) + \gamma(S_{1,t})\left(E(y_{t+k}) - y^*_{t+k}\right)\right] \\ &+ \rho \sum_{i=1}^{n} w_i i_{t-i} + u_t. \end{aligned} \quad (4)$$

As it is common in the literature,[6] it is assumed that the long-run inflation target is constant during the sample period so that it can be subsumed into the constant term in the regression, i.e.,

$$\alpha = \bar{r} - (\beta - 1)\pi^*.$$

Without further assumptions, values for both the long-run inflation target or the equilibrium real interest rate cannot be recovered simultaneously from the estimation. If one makes an assumption on either variable—e.g., by setting the equilibrium real interest rate equal to the ex-post sample real interest rate—one can derive a value for the other variable.

The coefficients on inflation and the output gap are allowed to depend on the unobservable state of the economy $S_{1,t}$ with the transition probabilities p_1 and q_1, i.e., they can take a different value in each regime:

$$\beta(S_{1,t}) \in \{\beta_1, \beta_2\}, \quad \gamma(S_{1,t}) \in \{\gamma_1, \gamma_2\}, \quad \beta_1 > \beta_2.$$

The state variable, $S_{1,t}$, can be thought of as representing the monetary policy regime that prevails at date t. The regimes are normalized such that the first regime has a

[6] See, e.g. Clarida, Galí, and Gertler (2000), Clarida and Gertler (1997) or Judd and Rudebusch (1998).

high coefficient on inflation. Neither sign or magnitude of coefficients nor the timing of the switches between both are imposed. The only assumption is that there are two different regimes for β and γ, with the estimation procedure classifying each observation as belonging to either regime with a certain probability.

In addition to switching in the coefficients, the variance of the error term is allowed to switch between a high variance and a low variance state.

$$\begin{aligned} u_t &\sim N(0, \sigma^2(S_{2,t})) \\ \sigma^2(S_{2,t}) &\in \{\sigma_1^2, \sigma_2^2\}, \quad \sigma_1^2 < \sigma_2^2 \end{aligned}$$

The process for the variance, $S_{2,t}$, follows a Markov-switching process that is independent from the process governing the switching in the coefficients. While the process $S_{1,t}$ represents shifts in the monetary policy preferences, the state variable $S_{2,t}$ accounts for shocks to the interest rate that are not captured in the reaction function, such as influences from the exchange rate. Periods of high and low variance can occur independently from the regime for the coefficients. Though one could let the variances switch simultaneously with the coefficients, we want to avoid that periods of high volatility dominate the assignment of different coefficients to restrictive and less restrictive periods of monetary policy. For example, during exchange rate crises for the EMS countries interest rates showed times of high volatility that were not necessarily related to the monetary policy regime. A specification with two independent Markov processes thus is more flexible since each monetary policy regime can be associated with a low and a high variance.

In the empirical specification the constant and the coefficient on the lagged dependent variable are not allowed to switch. It is thus assumed that the autocorrelation in the interest rate is not regime dependent, meaning that the central bank has the same concern for interest rate smoothing in both regimes.[7] The constant, which comprises the equilibrium real rate and the long-run inflation target, does not depend on the monetary policy regime as in the long run the equilibrium real rate should be independent from monetary policy. If the equilibrium real rate is computed as the sample ex-post real interest rate, the long-run inflation target can be derived as

$$\pi^* = \frac{\bar{r} - \alpha}{\beta(S_{1,t}) - 1}.$$

Even if the constant, α, is regime-independent and the real rate is assumed to be equal across regimes, the implied inflation objective, $\bar{r}$, will be different for each

[7] We do not want the autoregressive coefficient, ρ, to influence the classification of observations into either monetary policy regime so that $S_{1,t}$ should only drive switches in β and γ but not in ρ. Introducing a third Markov process for switches in the ρ would result in over-parametrization of the model.

regime. Clarida, Galí, and Gertler (2000) find that for the United States the difference between monetary rules in different time periods is better captured by different β coefficients than by different constants. If $\beta < 1$, the path for inflation is unstable because the central bank does not raise interest rates enough to keep pace with inflation, and there exists no equilibrium to which inflation would return. In this case, the implied inflation objective, π^*, is not defined.

For Germany and the United Kingdom the monetary policy rule is estimated as set up in equation (4). Estimation of a Taylor rule presumes that the country has been able to optimize between a domestic output and inflation target. For Italy and France this may not have been the case during the longer part of the sample period considered. Both countries were members of the EMS and had to conduct their monetary policy under an exchange rate constraint, so that they have not been able to consider only domestic variables in setting monetary policy.[8] To allow for an influence from the EMS on domestic monetary policy therefore the German interest rate is included into the reaction function for France and Italy.[9] In accordance with the literature, Germany is regarded as the leading country in the EMS, while the other member countries had to conduct their monetary policy such that they kept the exchange rate to the German mark fixed.

The specification for France and Italy thus supposes three different states for the coefficients. While the first two states are identically to those under monetary policy independence, in the third state the interest rate is set according to the German interest rate. The model thus becomes as follows:

$$\begin{aligned} i_t &= (1-\rho)\left[\alpha + \beta(S_{1,t})E(\pi_{t+j}) + \gamma(S_{1,t})\left(E(y_{t+k}) - y^*_{t+k}\right) + \lambda(S_{1,t})i_t^G\right] \\ &+ \rho\sum_{i=1}^{n} w_i i_{t-i} + u_t. \end{aligned} \tag{5}$$

with i^G denoting the German interest rate. The Markov process governing the regime switching for the coefficients now is allowed to switch between three different states with

$$\begin{aligned} \beta(S_{1,t}) &= \{\beta_1, \beta_2, 0\}, \\ \gamma(S_{1,t}) &= \{\gamma_1, \gamma_2, 0\}, \\ \lambda(S_{1,t}) &= \{0, 0, \lambda\}. \end{aligned}$$

[8]The influence of the EMS membership on monetary policy in the U.K. is not investigated as the U.K. was only member from October 1990 to September 1992. This episode is too short compared to the sample period of 28 years to give meaningful results.

[9]See Clarida, Galí, and Gertler (1998), Wyplosz (1999), Dolado, María-Dolores, and Naveira (2000), Mihov (2001).

The third state thus presumes that domestic variables play no role in setting monetary policy. Instead, the domestic interest rate is set in relation to the interest rate in Germany.

For the United States it is well documented that the time of the Volcker disinflation from 1979 to 1982 creates problems in the estimation of a monetary policy rule (see e.g. Fair 2001). As this period is characterized by an exceptionally high volatility of the interest rate, it seems warranted to allow for three states in the variance process for the United States.

$$u_t \sim N(0, \sigma^2(S_{2,t})), \quad \sigma^2(S_{2,t}) \in \{\sigma_1^2, \sigma_2^2, \sigma_3^2\}, \quad \sigma_1^2 < \sigma_2^2 < \sigma_3^2.$$

The regimes are ordered according to their residual variance. Like in the model for Germany and the United Kingdom, the process governing the switching coefficients can take two different values as in equation (4).

In the case of three possible regimes either for the coefficients or for the variance, the transition probability matrix is no longer 4×4 but 6×6 with eight transition probabilities to be estimated in the case of independent switching processes.[10] To restrict the number of parameters, the transition probability matrix governing the switches between the three regimes is assumed to show equal probabilities for a change into either one of the two other regimes. The transition probability matrix for the three-state Markov process thus is

$$P = \begin{pmatrix} p & (1-p)/2 & (1-p)/2 \\ (1-q)/2 & q & (1-q)/2 \\ (1-r)/2 & (1-r)/2 & r \end{pmatrix}.$$

This means that from regime 1 the Markov process can pass with equal probability into regime 2 or 3, and so on. Instead of estimating eight transition probabilities, this restriction reduces the number of parameters to five.

2.2 Markov Switching

For ease of exposition we refer in the following to the two-state model with two independent switching processes. The extension to a three-state process is straightforward. $S_{1,t}$ and $S_{2,t}$ are assumed to evolve according to a two-state, first-order Markov process, i.e., the probability $\Pr[S_{j,t=i}|\Psi_{t-1}]$, $j = 1, 2$ of being in a particular state $i = 0, 1$ in period t only depends on the state prevailing in period $t-1$. The case of two independent Markov processes can be regarded as switching between four different states, i.e., two regimes for the coefficients of the monetary policy reaction

[10] Six transition probabilities are coming form the three-state Markov process, plus two transition probabilities for the independent, two-state Markov process.

function with two different variances for each regime. Instead of considering two independent processes, a single Markov-switching process S_t is defined such that

$$\begin{aligned} S_t = 1 \quad &\text{if} \quad S_{1,t} = 0 \quad \text{and} \quad S_{2,t} = 0, \\ S_t = 2 \quad &\text{if} \quad S_{1,t} = 0 \quad \text{and} \quad S_{2,t} = 1, \\ S_t = 3 \quad &\text{if} \quad S_{1,t} = 1 \quad \text{and} \quad S_{2,t} = 0, \\ S_t = 4 \quad &\text{if} \quad S_{1,t} = 1 \quad \text{and} \quad S_{2,t} = 1. \end{aligned}$$

In general, the transition matrix for a four-state Markov process is 4×4 and would have 12 free parameters. With two independent switching processes, however, the transition probabilities for $S_{1,t} = 1$, $S_{1,t} = 0$, $S_{2,t} = 1$, and $S_{2,t} = 0$ have to sum up to unity individually. The elements of the transition probability matrix for the four states therefore can be written as products of the transition probabilities for both processes that are denoted as follows:

$$\begin{aligned} p_1 &= \Pr(S_{1,t} = 1 | S_{1,t-1} = 1), \\ q_1 &= \Pr(S_{1,t} = 0 | S_{1,t-1} = 0), \\ p_2 &= \Pr(S_{2,t} = 1 | S_{2,t-1} = 1), \\ q_2 &= \Pr(S_{2,t} = 0 | S_{2,t-1} = 0). \end{aligned} \tag{6}$$

The series S_t, $t = 1, 2, ..., T$ provides information about the regime the economy is in at date t. To obtain the probability for being in the first monetary policy regime, $S_{1,t} = 0$, one has to add the probabilities for the regimes $S_t = 1$ and $S_t = 2$. For the probability of being in the low variance state, $S_{2,t} = 0$, one has to add $S_t = 1$ and $S_t = 3$. If S_t were known before estimating the model, one could apply a dummy variable approach and define β, γ and σ as

$$\begin{aligned} \beta &= S_{1,t}\beta_0 + (1 - S_{1,t})\beta_1, \\ \gamma &= S_{1,t}\gamma_0 + (1 - S_{1,t})\gamma_1, \\ \sigma^2 &= S_{2,t}\sigma_0^2 + (1 - S_{2,t})\sigma_1^2. \end{aligned}$$

This would correspond to the approach Dolado, María-Dolores, and Naveira (2000) follow by choosing the episodes with inflation above and below target, or output above and below trend. Splitting the sample into the office terms of the different chairmen, like Clarida, Galí and Gertler (2000) and Judd and Rudebusch (1998) do, also can be regarded as a special case of the dummy variable approach. In the Markov-switching approach, however, S_t is assumed to be not observed, and the assignment of the regimes is estimated from the data.

To estimate the model, one has to consider the joint distribution $f(i_t, S_t|\Psi_{t-1})$ of i_t and S_t, which can be factored into

$$f(i_t, S_t|\Psi_{t-1}) = f(i_t|S_t, \Psi_{t-1})f(S_t|\Psi_{t-1}),$$

with $f(i_t, S_t|\Psi_{t-1})$ being the conditional normal density function for the regime $S_t = j$,

$$f(i_t, S_t|\Psi_{t-1}) = \frac{1}{\sqrt{2\pi}\sigma(S_t)} exp\left\{\frac{-\varepsilon_t^2(S_t)}{2\sigma^2(S_t)}\right\}, \tag{7}$$

and Ψ_{t-1} denoting information at time $t-1$ (see Kim and Nelson 1999). The likelihood function is thus a weighted average of the density functions for the four regimes, the weights being the probability of being in each regime,

$$\ln L = \sum_{t=1}^{T} \ln\left\{\sum_{j=1}^{4} f(i_t|S_t, \Psi_{t-1})\Pr\left[S_t = j|\Psi_{t-1}\right]\right\}.$$

Given a process for the evolution of the states, the model can be estimated using an iterative Maximum Likelihood procedure. $\Pr\left[S_t = j|\Psi_{t-1}\right]$ denotes the conditional probability that the t^{th} observation is generated by regime j. At the beginning of time t the probabilities are calculated as

$$\Pr\left[S_t = j|\Psi_{t-1}\right] = \sum_{k=1}^{4} \Pr\left[S_t = j|S_{t-1} = k\right]\Pr\left[S_{t-1} = k|\Psi_{t-1}\right],$$

where $\Pr\left[S_t = j|S_{t-1} = k\right]$ are the elements in the transition matrix that can be expressed in terms of p_1, q_1, p_2 and q_2, see equation (6). At the end of each period, the probabilities are updated using the following iterative filter (Kim and Nelson 1999),

$$\begin{aligned} \Pr\left[S_t = j|\Psi_t\right] &= \Pr\left(S_t = j|\Psi_{t-1}, i_t\right) \\ &= \frac{f(i_t|S_t = j, \Psi_{t-1})\Pr[S_t = j|\Psi_{t-1}]}{\sum_{j=1}^{4} f(i_t|S_t = j, \Psi_{t-1})\Pr\left[S_t = j|\Psi_{t-1}\right]}, \end{aligned} \tag{8}$$

with $f(i_t|S_t = i, \Psi_{t-1})$ as defined in equation (7). At the end of period t, the t^{th} observation of the dependent variable, i_t, contains new information about the state of the economy, S_t, in period t. Equation (8) shows that the conditional distribution of the state S_t, $\Pr\left[S_t = j|\Psi_{t-1}, i_t\right]$, is given by the conditional joint density of i_t and S_t, divided by the density of i_t. Once the model is estimated and $\Pr\left[S_t = j|\Psi_t\right]$ is generated, one can use an algorithm developed by Kim (1994) to estimate the smoothed probability for regime S_t using all information in the sample, i.e., $\Pr\left[S_t = j|\Psi_T\right]$, where $t = 1, 2, \ldots, T$.

The Markov-switching model is estimated using the BFGS algorithm in a recursive, non-linear optimization routine. All estimations are performed with RATS 5.0.

Starting values for the optimization routine are obtained from an OLS regression with a grid search over a plausible range of the switching coefficients to ensure that a global maximum of the likelihood function is attained.

3 The Data

The model is estimated with monthly data, starting in January 1973 for all countries. For France, Germany, and Italy as members of the European Monetary Union the sample ends in December 1998, because with January 1999 the responsibility for monetary policy went to the European Central Bank and a country-specific short-term interest rate does no longer exist. For the United States and the United Kingdom the sample period runs until December 2000.

Since the Taylor rule considers the interest rate as the monetary policy instrument the overnight money market rate is used. The output gap is calculated as the difference between the logarithm of industrial production and its trend value, which is computed with the HP-filter.[11] Figure 1 shows the short-term interest rate, inflation and the output gap. For all countries inflation and interest rates decline markedly during the 1990s. The exception is Germany where Unification lead to a rise in the interest rate and in inflation at the beginning of the 1990s. The volatility of all variables generally is higher at the beginning of the sample, which should be captured by the switching variance process. From 1973 to 1980 the ex-post real interest rate is low or even negative, and volatile. Thereafter it is consistently positive and has an average value between 3.5 % for Germany and the United States, and 5.7 % for Italy. The data already indicate that monetary policy might have changed in the course of time. Moreover, the countries show broad similarities in the development of the regression variables so that one can expect also to find similar monetary policy rules.

For the empirical analysis of equation (4) a measure of expected inflation and the expected output gap has to be constructed. For both variables dynamic forecasts from a vector autoregression (VAR) are used.[12] The VAR is specified with 12 lags of the interest rate, inflation and the output gap as endogenous variables. As exogenous variables a constant, the 12$^{\text{th}}$ lag of the slope of the yield curve, and the 12$^{\text{th}}$ lag

[11]For a detailed list of the data and the sources, see Appendix A.

[12]Expectations thus are not rational as expected inflation—which depends on past interest rates—is determined dynamically by using interest-rate forecasts from the VAR and not from the estimated switching Taylor rule. A joint estimation of the Taylor rule and expected inflation, however, is difficult because future inflation depends on the monetary policy regime so that expectations will be path dependent. As VARs are widely used in macroeconomic forecasting and the system includes the same endogenous variables as the monetary policy rule, the assumption that economic agents form their expectations according to this VAR seems justifiable.

of world commodity price inflation are included, so that forecasts are based only on variables dated time $t-1$ or earlier.[13] Expected inflation is computed as the twelve-step ahead forecast for inflation, and the one step-ahead forecast for the output gap is used. In other words, values of $j = 12$ and $k = 0$ are chosen for expectational terms in equation (4). Robustness of the estimation results to changes in these variables are examined in Section 4.2. There is no distinction between different regimes in this first-stage estimation. In the second step the monetary policy reaction function is estimated using maximum likelihood estimation as described above.

To account for autocorrelation in observed interest rates, equation (4) includes lags of the dependent variable. The number of lags is determined by a Lagrange-Multiplier (LM) test for first-order autocorrelation in a linear regression without Markov switching. For France, Germany, and the United States three lags of the dependent variable are included. For the other countries two lags seem sufficient to remove first-order autocorrelation from the residuals. Then, the switching model is estimated with the preferred number of lags. Results for the Markov-switching model are not sensitive to the lag length used.

4 Results for Time-Varying Taylor Rules

Table 1 gives the results for the estimation of equations (4) and (5).[14] The upper part of the table shows the coefficients for the first and the second regime. The constant and the autoregressive term are not allowed to switch with the monetary policy regime. For the autoregressive coefficients only the sum of the coefficients on the lagged interest rate is reported. T-values are calculated from an approximation to the inverse of the information matrix.

For all countries two distinct monetary policy regimes are found. In addition, for France and Italy a third regime emerges in which the German interest rate plays a role for domestic monetary policy. The coefficient on the expected inflation rate for the first regime is significant and greater than unity for all countries, ranging from 1.20 for Italy to 1.81 for the United States and implying an aggressive reaction towards expected inflation. In the second regime inflation coefficients for all countries are smaller than unity and are insignificant for Germany and Italy. This means that in the second regime the central bank adjusts the nominal interest rate

[13] As the coefficients are estimated over the whole sample period, forecast are in fact in-sample predictions. Due to the high number of parameters in a VAR, a recursive estimate would lead to a too high loss in degrees of freedom and unreliable forecasts especially at the beginning of the sample period.

[14] The reported coefficients are the long-run coefficients without the influence of the autoregressive terms, i.e., they give the values of α, β, γ and λ.

if inflation deviates from target, but not enough to drive also the real rate into the same direction. For France and Italy the German interest rate appears with a highly significant coefficient in the Taylor rule. For France, the coefficient estimate is with 0.82 close to the weight of 0.87 Dolado, María-Dolores, and Naveira (2000) find for the German interest rate in the French reaction function. While the weight for France is slightly below unity, the coefficient for Italy exceeds unity, implying that during the third regime Italy had to move its interest rate even more than Germany did. This results for Italy seems more plausible than the coefficients of -0.29 and -0.23 for the German interest rate in the Italian reaction function found by Mihov (2001) and Wyplosz (1999).

The output coefficients are generally less precisely estimated. For France, Germany, and the United States at least one output coefficient is significant on the 5% level, while for Italy both coefficients are significant only on the 10% level. As conjectured, the first regime—with a high weight on inflation—is associated with a lower coefficient on output than the second regime for France and the United States.This means that the monetary policy switches between a "dove" regime with a high weight on output and an anti-inflationary "hawk" regime. For Italy and the United Kingdom, output coefficients are insignificant so that monetary policy is mainly defined via inflation. For Germany, the regime with the low weight on inflation is also associated with a low weight on output, though point estimates for the output gap coefficient are relatively close together, meaning that the difference in the reaction to output is small between both regimes.[15]

The sum of the coefficients on the lagged dependent variable is between 0.91 and 0.96, which implies a high degree of persistence in interest rates. The reported t-statistics on the ρ coefficients are computed for the null hypothesis of $\rho = 1$. They indicate that ρ is significantly different from unity, so that the regressions do not contain a unit root.

The lower part of Table 1 gives the estimates for the standard deviation of the residuals in each regime as well as the transition probabilities p_1, q_1 and r_1 for switching in the monetary policy regime, and p_2, q_2 and r_2 for the variance switching process. For all countries one can clearly distinguish between a regime with a low residual standard deviation and one with a high standard deviation. The low standard deviation ranges from 0.17 for the United States to 0.30 for Italy; the high standard deviation is 5 to 10 times higher and lies between 1.13 for France and 2.03 for Germany.

All transition probabilities are close to unity, meaning that the regimes show

[15]The relative magnitude of the output coefficients for Germany and Italy is affected by the measure of the output gap, see Table 4.

high persistence—a feature that is common in the estimation of Markov switching models (see e.g., Sims 1999).

Figures 2 to 6 show the smoothed probabilities for the Markov processes governing the coefficient and the variance switching (on the left scale), together with the short-term interest rate (on the right scale).[16] At the beginning of the sample period, all countries except for Germany are in the second regime, which is characterized by an accommodative reaction to inflation. Before the second oil-price shock Germany also shortly switches into the accommodative regime, but is back in the aggressive regime after the oil-price shock at the time inflation starts to rise. After the inception of the EMS in 1979, Italy and France follow the German interest rate policy until around 1983. After 1983, France and Italy follow a conservative policy oriented towards domestic monetary variables. The United States follow a more conservative monetary policy starting at the end of the 1970s, while the United Kingdom switches to the aggressive regime only in 1983. Following the interest rate rise in Germany around German Unification, the German interest rate again became important for France, but not for Italy. With the recession of 1990/91 monetary policy in the United States became accommodative. For the European countries a switch into the accommodative regime happened after the EMS crisis in 1993. As at that time inflation was falling, an accommodative monetary policy in fact means that interest rates were not lowered enough to prevent the real interest rate from rising. Nominal interest rates fell, nevertheless, because of the central bank's reaction to the drop in the output gap. Only the United Kingdom, where the sample period is two years longer, experiences after a moderate increase in interest rates another transition to the first regime in 1997, when the Bank of England gained its independence.

Especially for Germany, a close correspondence between changes in interest rates and regime switches emerges. The first regime with a high weight on inflation is associated with periods of rising interest rates, e.g., from 1979 to 1982, and from 1988 to 1994. This means that the Bundesbank in such instances rises the interest rate more than inflation, and lets the real interest rate increase by 0.6 percentage points. In periods of falling interest rates—except for the years 1973 to 1977 and around 1993—the second regime prevails. In such situations interest rates fall only because of the output gap since in the second regime the estimate of the inflation coefficient, β_2, for the Bundesbank is close to zero. For the Bundesbank thus a bias towards a restrictive monetary policy is present, as rising inflation is countered more

[16]The interest rate as the endogenous variable in the estimation is depicted to ease interpretation of the timing of the states. As the monetary policy regimes depend on inflation and the output gap, the presentation of only one of these variables might be misleading. Nevertheless, inflation in general follows quite closely the movements of the interest rate, see Figure 1.

aggressively than falling inflation. This confirms the results by Dolado et al. (2000) who find that the Bundesbank raises the interest rate by 1.8 percentage points if inflation rises by 1 percentage point, but lets the real rate rise by 0.7 percentage points if inflation falls by 1 percentage point.

Figures 2 to 6 also show the smoothed probability for being in the low variance regime together with the short-term interest rate. While the model for the European countries allows switching between two different variances, for the United States three variance regimes are allowed. For the United States, the high-variance regime 3 is assigned solely to the time of the Volcker disinflation and the targeting of the monetary base from 1979 to 1983. The exceptionality of this episode is well documented in the literature.[17] Regime 1, which has the lowest variance, occurs from 1975 to 1979, at two occasions during the 1980s and during the 1990s. Regime 2 with a with an intermediate variance prevails until 1975 and during two short episodes in the 1980s. Though Germany shows the highest volatility, this state occurs only for a short time, mainly at the beginning of the sample before the Bundesbank adopted its strategy of monetary targeting. Concerning the timing of the variance switches, the European contries fall into two groups. For Germany and the United Kingdom, the high-variance regime concentrates at the beginning of the sample. While for Germany the high-variance state is confined to a relatively short period around the two oil-price shocks, the United Kingdom shows a much longer time of high volatility, which coincides with the failed attempts to bring inflation down at the beginning of the 1980s (Minford 1993). For France and Italy, variance regimes tend to be less persistent and switching is possibly caused by tensions in the EMS. This conjecture is supported by the positive correlation of 0.30 for Italy and 0.19 for France between the probability of being in the second regime and the annual depreciation of the Italian lira and the French franc vis-à-vis the German mark.

Table 2 shows the expected duration and the unconditional probabilities for the monetary policy and the variance switching regimes.[18] For France, Germany, and the United Kingdom expected duration of both monetary policy regimes ranges from one year to approximately four years. For Italy and the United States the expected duration of the monetary policy regimes is somewhat longer. The variance regimes last less than two years for most countries, except for the low variance regime in the United States, which has a expected duration of around six years. Except for France, the unconditional probability of being in the first, anti-inflationary regime

[17]See Fair (2001), Clarida, Galí, and Gertler (2000), Sims (1999), or Judd and Rudebusch (1998).

[18]The unconditional probabilities were computed according to the formula given in Garcia and Perron (1996, p. 113).

is between 30% and 40%. The unconditional probability of the low variance regime lies around 80% for France, Germany, and Italy. In the United Kingdom the low and the high variance regime have an unconditional probability of 50% each, which reflects the much longer time span of high volatility for U.K. compared to the other countries. In the United States the low variance state occurs with an unconditional probability of 63%.

4.1 Comparison to a Linear Model

As a Markov-switching model permits more flexibility than a linear model, a comparison between both types of models is of interest. Figure 7 compares the distribution of the standardized residuals from the Markov-switching specification to the residuals from a linear regression model containing the same variables. It is apparent that the Markov-switching model improves the distribution of the residuals, especially by reducing the number of outliers, i.e., by lowering excess kurtosis as compared to the linear model. This effect is primarily achieved by incorporating switching between a low and a high residual variance.

The top panel of Table 3 shows skewness and excess kurtosis for the residuals from the Markov-switching model and the linear model. For all countries both measures are much closer to the theoretical values of a normally distributed variable in the Markov-switching model than in the linear model. Also the value of the log likelihood function increases considerably for the switching specification as compared to a linear model. The test statistic for a likelihood ratio test of the Markov switching against the linear model would lie above 200 for all countries. Though likelihood ratio tests in this case are not strictly applicable as some nuisance parameters are not identified under the null hypothesis (see e.g., Garcia 1998), the size of the test statistic indicates that the Markov switching model is indeed superior to the linear model.

The effect of switching coefficients for inflation and the output gap can be visualized by looking at the implied target interest rate. A monetary policy rule is only a shortcut to describe central-bank behavior since actually central banks take more information into account than can be reflected in a simple rule. A comparison of actual interest rates to the implied target rates from the monetary rule therefore can give an indication about the appropriateness of the rule to reflect actual central bank behavior.

Figure 8 shows the implied target rates from the Markov-switching model and the linear model together with the actual interest rate. For the computation of the implied target rate the autoregressive terms are excluded so that only the influence from the coefficients on output and inflation and the constant determine the implied target rate.[19] For France and Italy, the Markov-switching model improves over the linear model especially at the beginning of the sample period, where the linear specification shows large swings in the implied target rate. For the United Kingdom the Markov model captures the change in monetary policy that followed the entry into the EMS and the adoption of inflation targeting two years later, while the linear specification does not. At the beginning of the sample period, however, the implied rates for both models are persistently higher than the actual rate. The bottom panel of Table 3 shows the mean squared error (MSE) for the deviation of the actual interest rate from the implied target rate, again without considering the autoregressive coefficients. Note that this is not a test of goodness of fit for the Markov-switching versus the linear model but only a comparison of the closeness of the implied interest rate to the actually observed interest rate. For all countries the Markov-switching model constitutes a clear improvement, but especially so for France and Italy.

4.2 Robustness

Finally, robustness of the results with respect to changes in the definition of expected inflation and the output gap is checked. First, instead of using the HP-filter the output gap is computed as the deviation of the logarithm of industrial production from a linear and quadratic trend, see Clarida, Galí, and Gertler (1998). Second, actual inflation and the actual HP-filtered output gap are used instead of the forecasted variables.

Table 4 shows the results for the model with expected inflation and the expected output gap, computed as the deviation from a linear and a quadratic trend. Table 5 gives the results for the estimation with actual inflation and the actual HP-filtered output gap. For brevity, both tables only report the economically interesting β and γ coefficients. Table 4 shows that size and significance of the inflation coefficients remain basically unchanged when the output gap is computed with a linear and a quadratic trend. For all countries except for the United Kingdom the second regime now implies a stronger reaction to output than the first regime, thus confirming a switch between a "hawk" regime with a high weight on inflation and a low one on the output gap, and a "dove" regime with reversed weights. In Table 5 inflation coefficients turn out to be somewhat lower than in the benchmark case. In contrast to the other specifications, the output response in the second regime now is significant for the United Kingdom. With the specification in Table 5 now only France, the

[19] With the autoregressive terms the fitted values would be much closer for both models so that the difference between actual and fitted interest rates would be hardly discernible in the graph.

United Kingdom, and the United States show output responses that corresponds to notion of "hawk" and "dove" regimes. In general, however, results are robust to the change in specification.

5 Conclusion

This paper adds to the growing literature on time-varying monetary policy rules. Switching models are an interesting alternative to a conventional linear specification of a monetary policy rule, as the effects of a changing economic environment on monetary policy can be investigated without having to determine the dates of the changes exogenously.

In this paper monetary policy reaction functions for France, Germany, Italy, the United Kingdom, and the United States have been estimated, using a Markov-switching model with independent switching processes for the coefficients of the monetary policy regime and the residual variance. The results show that for all central banks the weights assigned to inflation and the output gap switch between different states. One regime is associated with a high weight on inflation, the other implies that the central bank follows an accommodative policy. Following Owyang and Ramey (2000) regimes can be classified as a "dove" regime with a high weight on output and a low weight on inflation, and a "hawk" regime with a high weight on inflation and a low one on output.

Switching in the residual variance apparently captures the effects of one-time shocks like supply shocks, changes in external constraints or pressure on the exchange rate that are not modelled explicitly in the monetary policy rule. As compared to a linear model containing the same variables, the Markov switching model improves the distribution of the residuals as well as the performance of implied interest rates.

A Appendix: Data

The interest rate is the money market rate from the International Financial Statistics (IFS) CD-Rom of the International Monetary Fund for the European countries and the federal funds rate for the United States. All data for the United States are from the database of the Federal Reserve Bank of St. Louis (FRED). Inflation is measured by the annual change in the consumer price index (CPI). The output gap is calculated as the difference between the logarithm of industrial production and its trend value, which is obtained by the HP-filter with the usual weight of 14400 for monthly data. Since it is well known that the HP-filter might give unreliable trend estimates at the end of the sample period, the filter was run over a period from the first quarter of 1970 to the last quarter of 2000 for all countries. This longer sample period was also used to compute the linear and quadratic trend estimates later in the paper.

Industrial production and CPI data are from the Main Economic Indicators of the OECD, except for Germany where the CPI is from the Monthly Reports of the Deutsche Bundesbank. The CPI relates to West Germany only because we want to avoid to deal with the breaks in the data caused by German Unification, which was followed by a successive lifting of price controls in East Germany. Due to the limited economic size of Eastern Germany the West German inflation rate differs only marginally from inflation for total Germany. West German industrial production is linked to the series for unified Germany after 1991, see OECD (2000).

References

Amano, R.A., D. Coletti, and T. Macklem (1999): "Monetary Rules when Economic Behaviour Changes," in *Hunt, B. and A. Orr (eds.)*: Monetary Policy under Uncertainty: Workshop Held at the Reserve Bank of New Zealand 29-30 June 1998, Reserve Bank of New Zealand, Wellington, pp. 157-200.

Batini, N. and A.G. Haldane (1999): "Forward-Looking Rules for Monetary Policy," in *Taylor, J. B. (ed.)*: Monetary Policy Rules, National Bureau of Economic Research, Chicago, pp. 157-192.

Blinder, A.S. (1998): "Central Banking in Theory and Practice," MIT Press, Cambridge, Mass.

Christiano, L.J. and C.J. Gust (1999): "Taylor Rules in a Limited Participation Model," *De Economist* 147 (4), pp. 437-460.

Clarida, R., J. Galí, and M. Gertler (1998): "Monetary Policy Rules in Practice: Some International Evidence," *European Economic Review* 42 (6), pp. 1033-1068.

Clarida, R., J. Galí, and M. Gertler (2000): "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory," *Quarterly Journal of Economics* 115 (1), pp. 147-180.

Clarida, R. and M. Gertler (1997): "How the Bundesbank Conducts Monetary Policy," in *Romer, C. D. and D. H. Romer (eds.)*: Reducing Inflation: Motivation and Strategy, University of Chicago Press, Chicago, pp. 363-412.

Cukierman, A. (1999): "The Inflation Bias Result Revisited," *Foerder Institute for Economic Research, Tel-Aviv University, Working Paper,* No. 99-38.

Deutsche Bundesbank (1999): "Taylor-Zins und Monetary Conditions Index," *Monatsberichte* April, pp. 47-63.

Dolado, J. J., R. María-Dolores, and M. Naveira (2000): "Asymmetries in Monetary Policy - Evidence for Four Central Banks," *CEPR Discussion Paper Series,* No. 2441.

Dueker, M.J. and A. Fischer (1996): "Inflation Targeting in a Small Open Economy: Empirical Results for Switzerland," *Journal of Monetary Economics* 37, pp. 89-103.

Engel, C.M. and J.D. Hamilton (1990): "Long Swings in the Dollar: Are They in the Data and Do Markets Know it?," *American Economic Review* 80 (4), pp. 689-713.

Fair, R.C. (2001): "Actual Federal Reserve Policy Behavior and Interest Rate Rules," *Federal Reserve Bank of New York, Economic Policy Review* 7 (1), pp. 61-72.

Fischer, S. (1994): "Modern Central Banking," in *Capie, F. and others (eds.)*: The Future of Central Banking, Cambridge University Press, Cambridge, pp. 262-308.

Garcia, R. (1998): "Asymptotic Null Distribution of the Likelihood Ratio Test in Markov Switching Models," *International Economic Review* 39, pp. 763-788.

Garcia, R. and P. Perron (1996): "An Analysis of the Real Interest Rate under Regime Shifts ," *Review of Economics and Statistics* 78 (1), pp. 111-125.

Gerlach, S. (2000): "Asymmetric Policy Reactions and Inflation," *unpublished manuscript, Bank for International Settlements.*

Hamilton, J.D. (1989): "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle," *Econometrica* 57 (2), pp. 357-384.

Jordan, T.J. (2001): "Monetary Control Uncertainty and Inflation Bias," *Journal of Economics* 73 (2), pp. 125-147.

Judd, J.P. and G.D. Rudebusch (1998): "Taylor's Rule and the Fed: 1970-1997," *Federal Reserve Bank of San Francisco, Economic Review* (3), pp. 3-16.

Kim, C.-J. (1994): "Dynamic Linear Models with Markov-Switching," *Journal of Econometrics* 60, pp. 1-22.

Kim, C.-J. and C.R. Nelson (1999): "State-Space Models with Regime Switching: Classical and Gibbs-Sampling Approaches with Applications," MIT Press, Cambridge, Mass.

Kozicki, S. (1999): "How Useful Are Taylor Rules for Monetary Policy?," *Federal Reserve Bank of Kansas City, Economic Review* 84 (2), pp. 5-34.

Levin, A., V. Wieland, and J.C. Williams (1999): "Robustness of Simple Monetary Policy Rules under Model Uncertainty," in *Taylor, J. B. (ed.)*: Monetary Policy Rules, National Bureau of Economic Research, Chicago, pp. 263-299.

Mihov, I. (2001): "Monetary Policy Implementation and Transmission in the European Monetary Union," *Economic Policy* 33, pp. 370-406.

Minford, P. (1993): "United Kingdom," in *Fratianni, M. U. and D. Salvatore (eds.)*: Monetary Policy in Developed Economies, Greenwood Press, Westport, CT, pp. 405-432.

Neumann, M. J. M. and J. von Hagen (2002): "Does Inflation Targeting Matter?," *Zentrum für Europäische Integrationsforschung, Bonn, Working Paper,* No. B01.

OECD (2000): "Main Economic Indicators, Sources and Definitions," *www.oecd.org /std/meimeta.pdf,*.

Owyang, M. and G. Ramey (2001): "Regime Switching and Monetary Policy Measurement," *Federal Reserve Bank of St. Louis, Working Paper,* No. 002A.

Romer, C.D. and D.H. Romer (1989): "Does Monetary Policy Matter? A New Test in the Spirit of Friedman and Schwartz," in *Blanchard, O. J. and S. Fischer (eds.)*: NBER Macroeconomics Annual, MIT Press, Cambridge, pp. 121-170.

Rudebusch, G.D. (2001): "Is the Fed Too Timid? Monetary Policy in an Uncertain World," *Review of Economics and Statistics* 83 (2), pp. 203-217.

Ruge-Murcia, F. J. (2001): "Inflation Targeting under Asymmetric Preferences," *Banco de España, Documento de Trabajo,* No. 0106.

Sims, C. A. (1999): "Drift and Breaks in Monetary Policy," *unpublished manuscript, Princeton University.*

Smets, F. (2002): "Output Gap Uncertainty: Does it Matter for the Taylor Rule?," *Empirical Economics* 27 (1), pp. 113-129.

Taylor, J.B. (1993): "Discretion versus Policy Rules in Practice," *Carnegie-Rochester Conference Series on Public Policy* 39, pp. 195-214.

Wyplosz, C. (1999): "Towards a More Perfect EMU," *CEPR Discussion Paper Series,* No. 2252.

Tables and Figures

Table 1: Results for Taylor Rule Estimation.

	France	*Germany*	*Italy*	*U.K.*	*U.S.*
α	2.96 (7.73)	3.09 (11.36)	5.34 (3.40)	2.42 (1.72)	2.19 (2.30)
β_1	1.24 (9.23)	1.55 (10.91)	1.20 (8.76)	1.70 (7.75)	1.81 (4.79)
β_2	0.45 (5.92)	-0.01 (-0.04)	0.23 (1.55)	0.74 (2.43)	0.81 (3.56)
γ_1	0.02 (0.05)	0.92 (4.52)	0.81 (1.71)	0.30 (0.58)	-0.37 (-1.26)
γ_2	0.43 (3.84)	0.66 (2.48)	0.55 (1.73)	0.27 (0.47)	1.01 (3.39)
λ	0.82 (6.63)		1.40 (5.93)		
ρ	0.91 (4.30)	0.92 (2.88)	0.93 (2.52)	0.93 (2.82)	0.96 (4.08)
σ_1	0.19 (10.77)	0.18 (8.48)	0.30 (14.34)	0.22 (8.90)	0.17 (16.54)
σ_2	1.13 (5.77)	2.03 (5.15)	1.42 (5.89)	1.38 (9.39)	0.54 (8.45)
σ_3					1.73 (4.46)
p_1	0.93 (8.89)	0.96 (20.29)	0.99 (59.17)	0.96 (22.24)	0.99 (95.82)
q_1	0.96 (13.89)	0.96 (21.27)	0.99 (46.79)	0.98 (35.09)	0.99 (157.98)
r_1	0.93 (15.06)		0.98 (57.22)		
p_2	0.92 (57.60)	0.97 (36.92)	0.95 (65.10)	0.95 (25.55)	0.99 (135.80)
q_2	0.73 (8.47)	0.85 (6.88)	0.76 (8.46)	0.95 (19.38)	0.94 (29.27)
r_2					0.96 (13.08)
L	-143.82	-129.21	-210.67	-348.59	-70.81

Note: Estimation of equations (4) and (5) in the text. T-values in parentheses; the t-value for ρ is computed for the null hypothesis of $\rho = 1$. The sample period is 1973:1 to 1998:12 for the EMU countries and 1973:1 to 2000:12 for the United Kingdom and the United States. The transition probabilities p_1, q_1 and r_1 are associated with switching in the coefficients, p_2, q_2 and r_2 with variance switching. L is the value of the log likelihood function.

Table 2: Expected Duration and Unconditional Probability of Regimes.

	France	*Germany*	*Italy*	*U.K.*	*U.S.*
Mean duration in years					
Monetary policy regime 1	1.11	1.87	6.42	2.37	6.55
Monetary policy regime 2	2.07	2.22	7.22	3.72	14.43
Monetary policy regime 3	1.21		3.43		
Variance regime 1	1.06	2.51	1.58	1.72	5.90
Variance regime 2	0.31	0.56	0.35	1.74	1.45
Variance regime 3					1.99
Unconditional probabilities					
Monetary policy regime 1	0.25	0.46	0.38	0.39	0.31
Monetary policy regime 2	0.47	0.54	0.42	0.61	0.69
Monetary policy regime 3	0.28		0.20		
Variance regime 1	0.78	0.82	0.82	0.50	0.63
Variance regime 2	0.22	0.18	0.18	0.50	0.16
Variance regime 3					0.21

Note: Values derived from the estimates in Table 1.

Table 3: Comparison of Markov-Switching Model to Linear Model.

	France	*Germany*	*Italy*	*U.K.*	*U.S.*
Skewness MS model	0.50	−0.04	0.10	−0.08	−0.10
Skewness linear model	1.87	−0.41	1.38	−0.02	−1.98
Kurtosis MS model	0.69	1.85	0.28	1.29	0.92
Kurtosis linear model	11.12	19.07	11.89	4.90	28.72
L MS model	−143.82	−129.21	−210.67	−348.59	−70.81
L linear model	−273.36	−382.11	−315.08	−484.40	−308.39
MSE MS model	4.43	5.52	8.75	12.04	8.50
MSE linear model	6.08	6.24	26.48	13.38	11.54

Note: The first panel shows skewness and excess kurtosis (i.e. above the theoretical value of 3 for a normal distributed variable) for the residuals from the Markov-switching (MS) and the linear model. The second panel gives the value of the log likelihood function. The bottom panel shows the mean squared error for the deviation of the implied target interest rate from the actual interest rate.

Table 4: Taylor-Rule with Detrended Output.

	France	*Germany*	*Italy*	*U.K.*	*U.S.*
β_1	1.26 (13.42)	1.89 (2.70)	1.48 (5.72)	1.34 (2.63)	1.87 (6.10)
β_2	0.39 (5.00)	0.17 (0.51)	0.35 (1.18)	0.54 (1.25)	0.49 (2.34)
γ_1	-0.08 (-0.44)	0.16 (0.30)	0.36 (1.83)	0.31 (0.86)	-0.23 (-1.00)
γ_2	0.29 (3.65)	0.71 (5.48)	0.49 (1.60)	-0.04 (-0.08)	0.36 (3.52)
L	-142.74	-129.09	-211.80	-344.71	-68.44

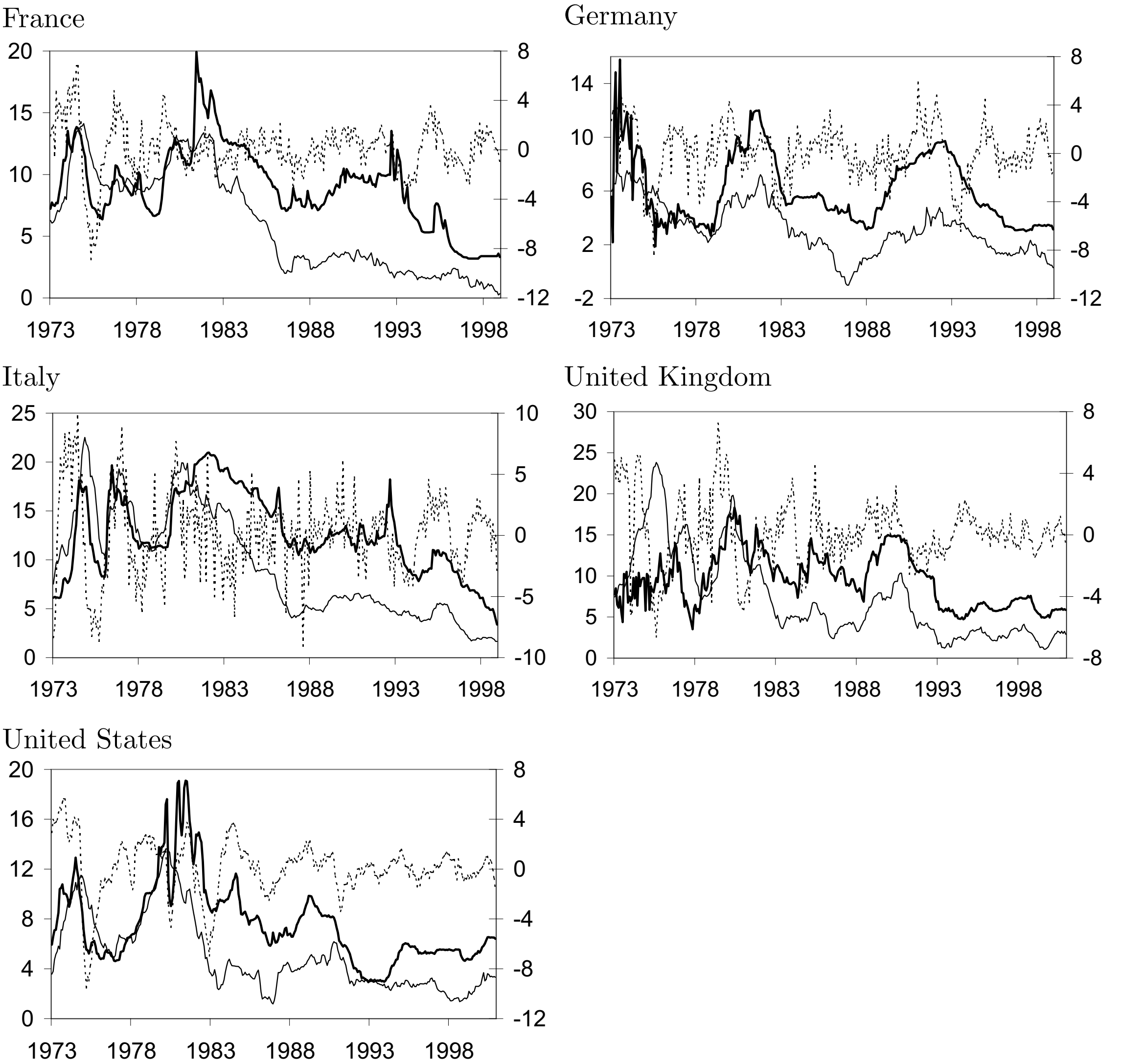
Note: The table gives the coefficients on inflation and output for the models in equations (4) and (5) with expected inflation 12 months ahead and the expected, contemporaneous deviation of output from a linear and quadratic trend.

Table 5: Taylor-Rule with Actual Inflation.

	France	*Germany*	*Italy*	*U.K.*	*U.S.*
β_1	1.02 (11.61)	1.53 (8.69)	1.04 (11.07)	1.10 (9.42)	1.46 (4.46)
β_2	0.44 (5.53)	-0.03 (-0.12)	0.31 (2.22)	-0.68 (-2.88)	0.02 (0.08)
γ_1	0.39 (0.37)	0.65 (3.38)	0.59 (2.11)	0.36 (0.90)	1.90 (3.06)
γ_2	0.41 (3.96)	0.27 (1.22)	0.26 (1.46)	1.40 (2.20)	2.65 (3.04)
L	-150.68	-142.97	-214.92	-342.95	-69.18

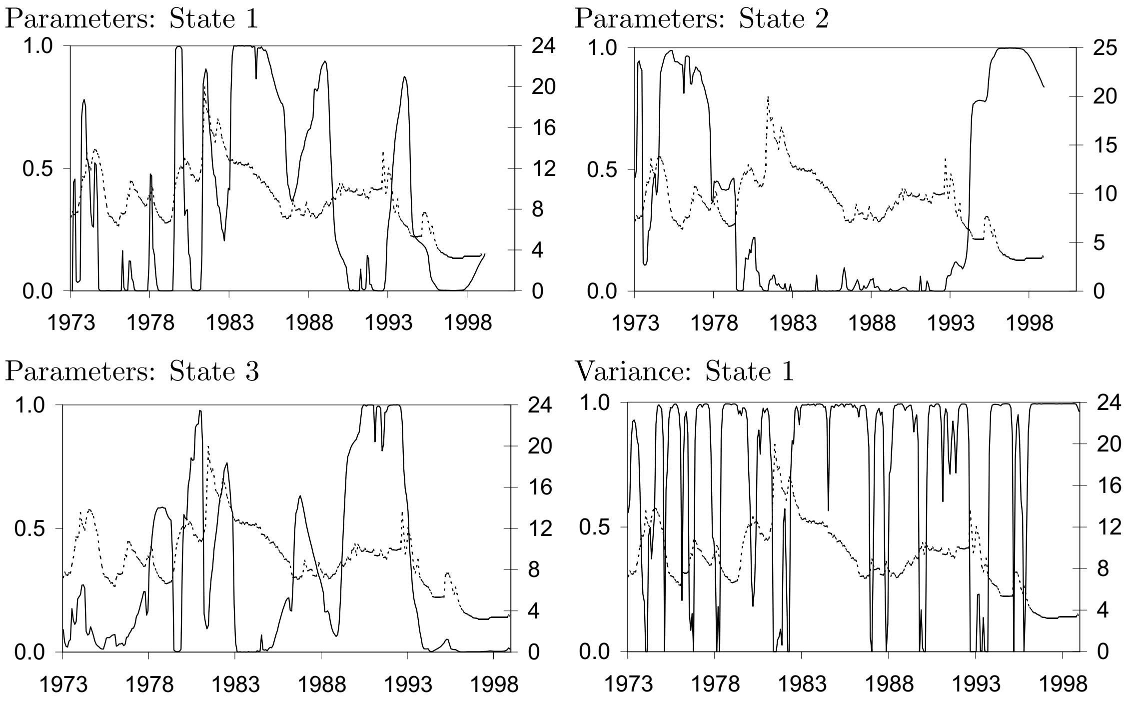
Note: The table gives the coefficients on inflation and output for the models in equations (4) and (5) with actual inflation and the actual output gap, computed as the deviation from a HP-filtered trend.

Figure 1: Interest Rate, Inflation and Output Gap.



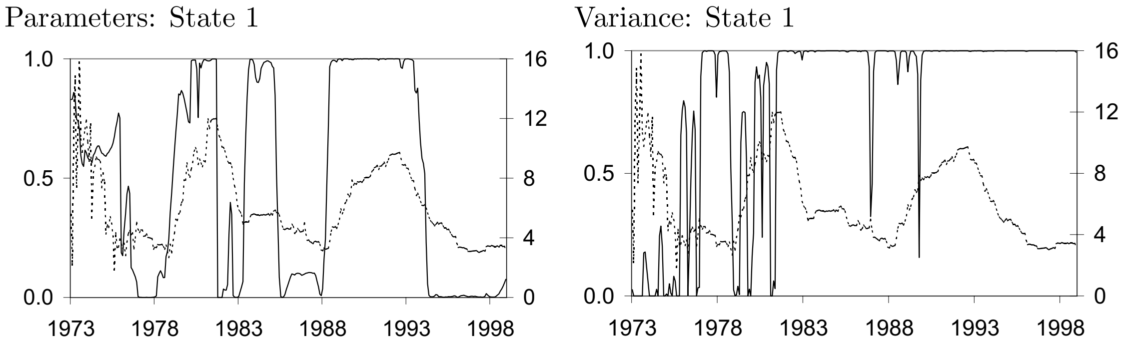
Note: The figure shows the short-term interest rate (thick line) and inflation (thin line) in percent on the left scale. The output gap (broken line) is shown on the right scale.

Figure 2: Smoothed Probabilities for France.



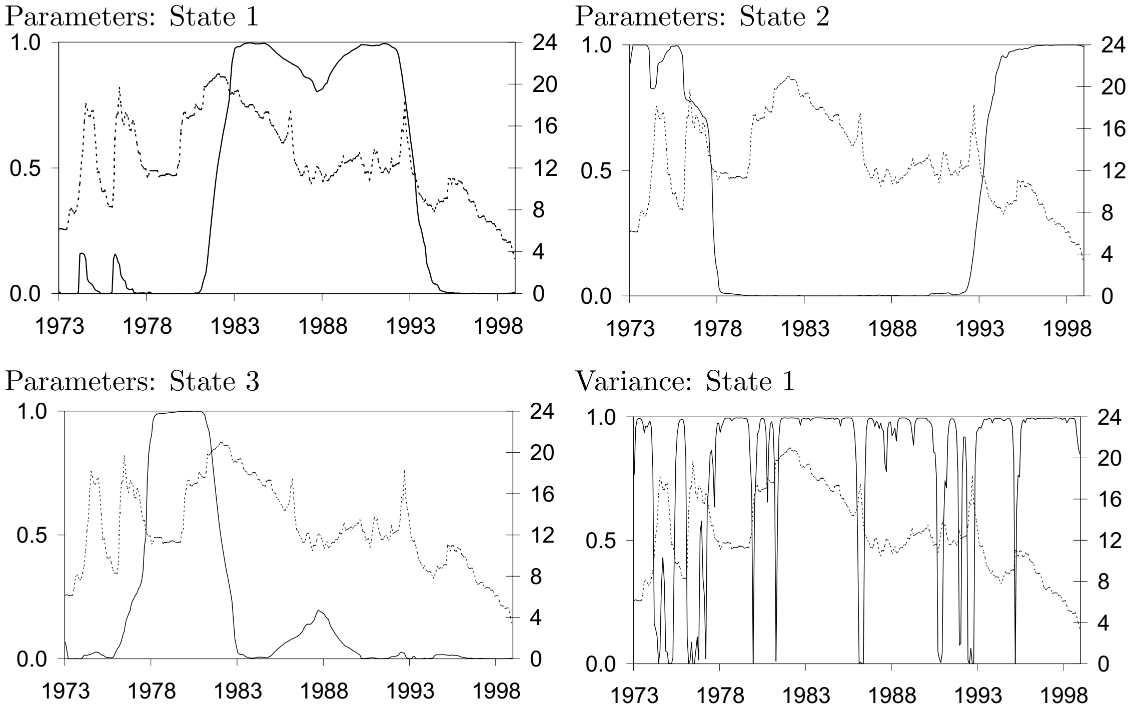
Note: The figure shows the smoothed probability from the model in Table 1 (straight line, left scale) together with the short-term interest rate (broken line, right scale).

Figure 3: Smoothed Probabilities for Germany.



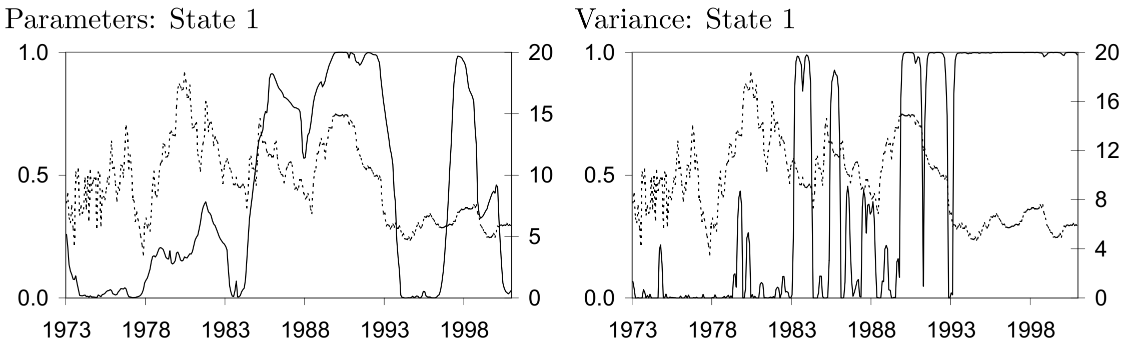
Note: The figure shows the smoothed probability from the model in Table 1 (straight line, left scale) together with the short-term interest rate (broken line, right scale).

Figure 4: Smoothed Probabilities for Italy.



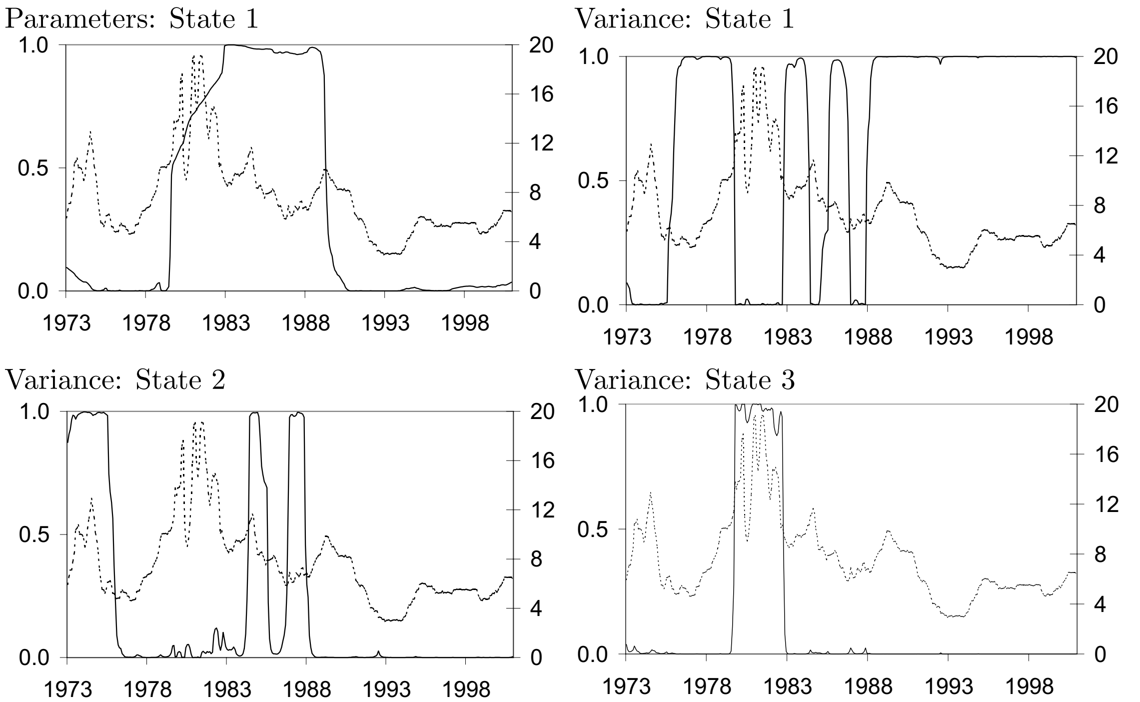
Note: The figure shows the smoothed probability from the model in Table 1 (straight line, left scale) together with the short-term interest rate (broken line, right scale).

Figure 5: Smoothed Probabilities for the United Kingdom.



Note: The figure shows the smoothed probability from the model in Table 1 (straight line, left scale) together with the short-term interest rate (broken line, right scale).

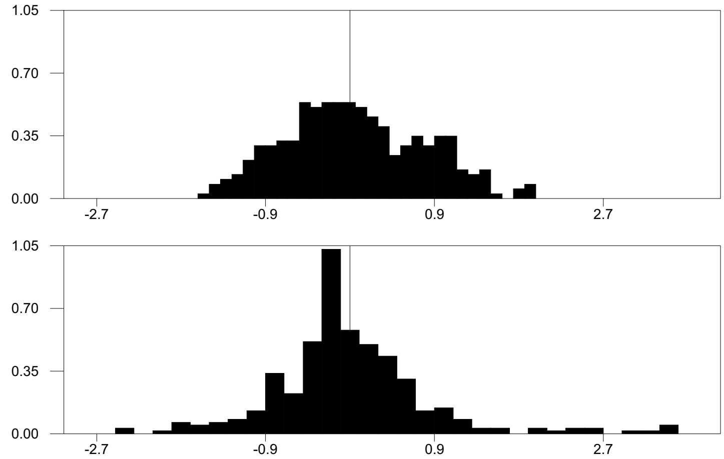
Figure 6: Smoothed Probabilities for the United States.



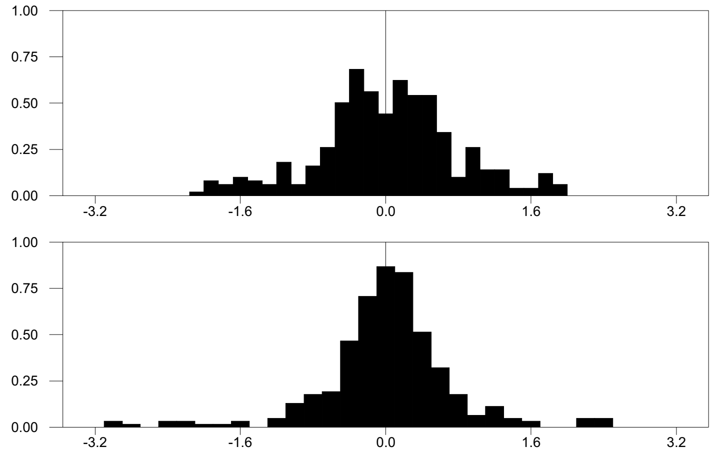
Note: The figure shows the smoothed probability from the model in Table 1 (straight line, left scale) together with the short-term interest rate (broken line, right scale).

Figure 7: Residuals from Markov-Switching Model and Linear Model.

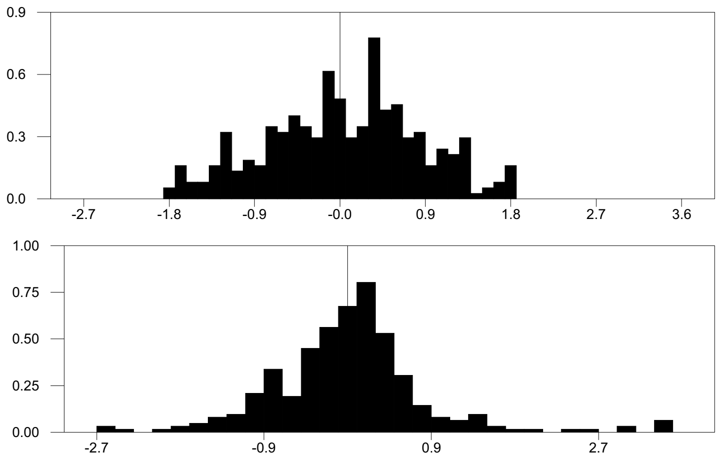
France



Germany



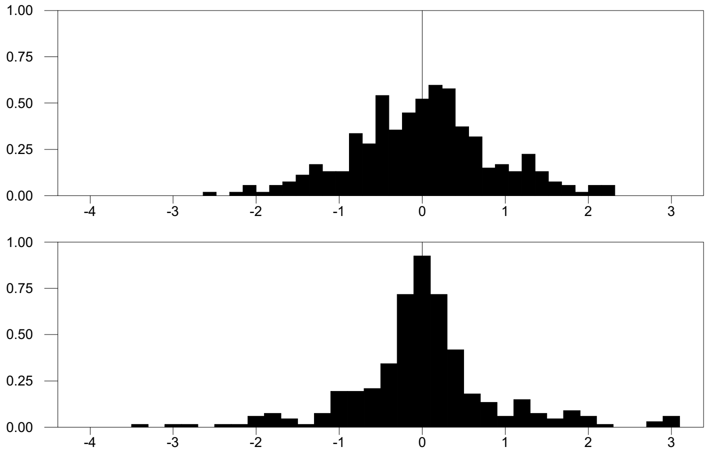
Italy



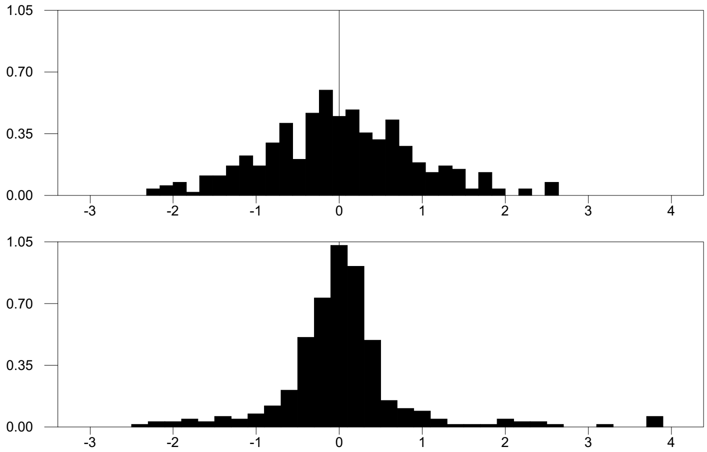
Note: The first panel presents the standardized residuals from the Markov-switching model in Table 1, the second panel the residuals from a linear regression on the same variables.

Figure 7 (cont.): Residuals from Markov-Switching Model and Linear Model.

United Kingdom

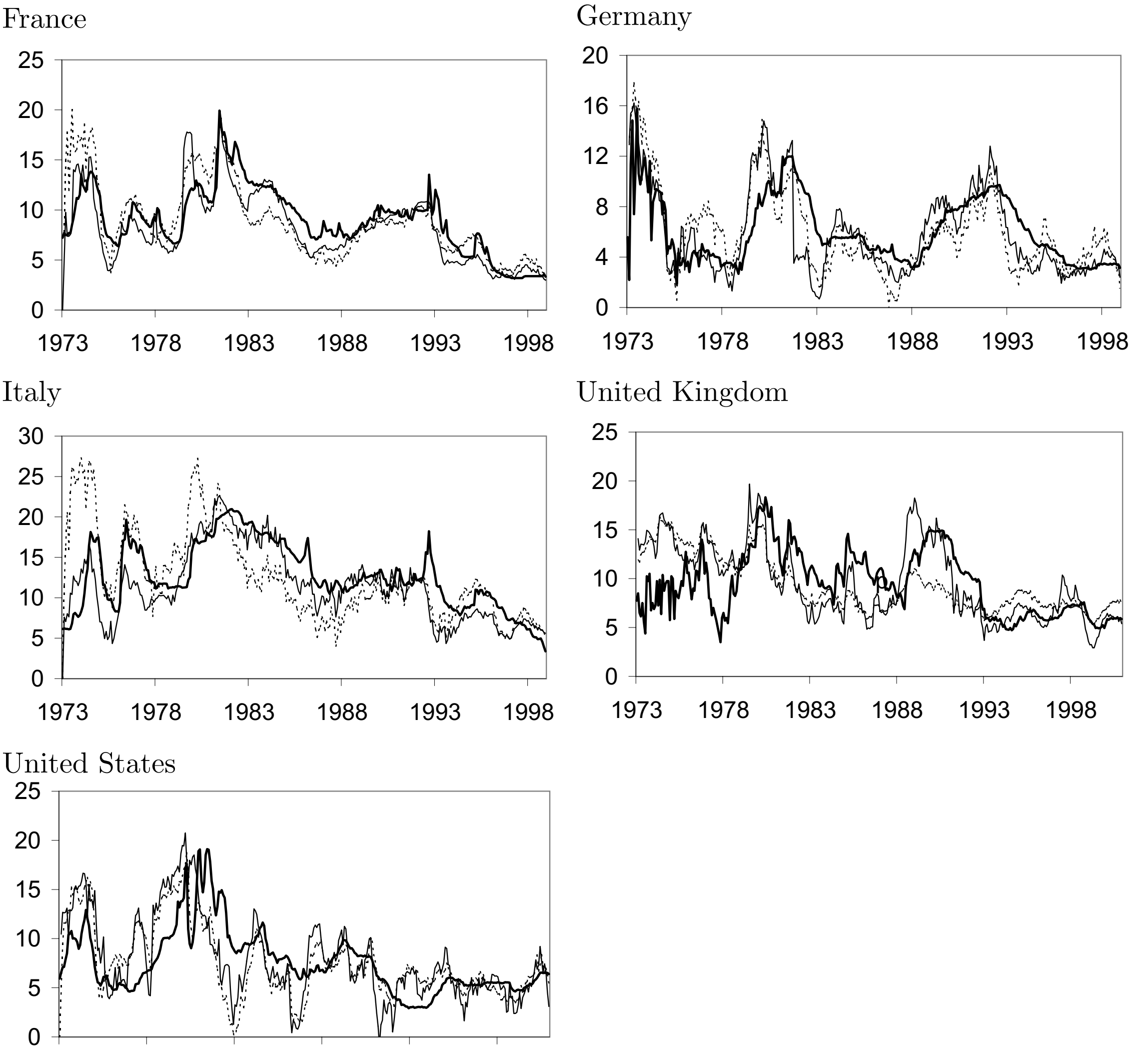


United States



Note: The first panel presents the standardized residuals from the Markov-switching model in Table 1, the second panel the residuals from a linear regression on the same variables.

Figure 8: Implied Target Interest Rate from Markov-Switching and Linear Model.



Note: The thick straight line is the actual short-term interest rate, the thin straight line the implied interest rate from the Markov-switching model and the thin broken line the implied interest rate from the linear model. Interest rates are in percent.