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A NOTE ON TIME VARIATION IN A FORWARD-LOOKING MONETARY POLICY RULE: EVIDENCE FROM EUROPEAN COUNTRIES

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This paper estimates time-varying forward-looking monetary policy reaction functions for the central banks of France, Germany, Italy, and the United Kingdom. We utilize the framework developed by Kim [*Economics Letters* 91 (2006) 21–26] and Kim and Nelson [*Journal of Monetary Economics* (2006) 1949–1966] to deal with the issue of endogeneity in a time varying–parameter model. Our results find substantial time variation in the conduct of monetary policy in these four countries, which cannot be adequately captured by the conventional fixed-coefficient approach. Our findings suggest that there was a significant decline in the Bank of France's and the Bank of Italy's response to the German interest rate in 1992, and it coincided with the breakdown of the exchange rate management system in Europe. Our results suggest that the Bank of England was slower than the Bundesbank to increase its response to expected inflation, as its response to inflation became more than one-for-one only in the early 1980s.

Keywords: Endogeneity, Forward-Looking Monetary Policy, Heteroskedasticity, Nonlinearity, Time-Varying Parameter Model

1. INTRODUCTION

The Taylor rule (1993) is widely used as a simple tool for assessing the monetary policy performance of central banks in different countries. In an influential paper, Clarida, Gali, and Gertler (2000) proposed a forward-looking version of the Taylor rule.[1] It shows how central banks respond to expected deviations of inflation and the output gap from their target levels. Using this reaction function, Clarida, Gali, and Gertler (1998; CGG hereafter) estimated a fixed-coefficient monetary policy reaction function for a set of European countries. They found that the Bundesbank had pursued implicit inflation targeting since 1979. Their findings suggested that the monetary policies of France and Italy were heavily influenced by the German

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monetary policy, and the interest rates in these countries were much higher than domestic macroeconomic conditions warranted.

Cogley and Sargent (2001) and Boivin (2006) argued that the fixed-coefficient approach to estimation of a monetary policy rule is not able to capture the time-varying response of the Federal Reserve to inflation and the output gap in the United States. Therefore, a time-varying parameter (TVP) model is needed to capture the gradual evolution of monetary policy. Possible sources of time variation include changing views of the central bank about the role of monetary policy, evolving institutional features of the central bank, relative importance of inflation stabilization versus real activity stabilization in different time periods, and different response to changes in inflation and output gap at different levels of inflation and output gap [Dolado et al. (2005)].[2] Boivin (2006) found that allowing for time variation in response to movements in inflation and expected gap provides a better understanding of the evolution of monetary policy in the United States, and it also reconciles the conflicting results that are obtained when different types of data[3] are used in a fixed-coefficient approach.

This paper estimates a time-varying version of the forward-looking monetary policy rule proposed by Cogley and Sargent (2001) and Boivin (2006) and applies it to four major European countries: France, Germany, Italy, and the United Kingdom. Allowing the central bank's conduct of monetary policy to vary with time is especially important in these countries, because of external constraints such as the breakdown of the Bretton Woods system in the 1970s, the formation of the European Monetary System (EMS), and the adoption of explicit inflation targeting in countries such as the United Kingdom. Assenmacher-Wesche (2006) also attempts to capture the time variation in monetary policy formulation in Germany and the United Kingdom. using a Markov switching model that allows shifts in the central banks' reaction coefficients as well as residual variances. She finds that the Bundesbank attached a higher relative weight to inflation than the Fed. However, Assenmacher-Wesche's paper does not use a forward-looking version of the Taylor rule because of the endogeneity of the regressors (future inflation and future output gap). This endogeneity problem makes estimation in a conventional regime-switching model infeasible.

The extension of a fixed-coefficient forward-looking monetary policy reaction function to a time-varying framework is not straightforward. The regressors in the forward-looking reaction function (future inflation and the future output gap) are correlated with the error term, thus making OLS estimates inconsistent.[4] Conventional IV or conventional GMM estimation procedures cannot be applied directly to a time-varying parameter model when the regressors are endogenous. Using the methodology introduced in Kim (2006) and Kim and Nelson (2006),[5] we are able to deal with time variation in the response coefficients and endogeneity to produce consistent estimates. In addition, this approach allows us to handle the issue of nonlinearity that arises due to interest-rate smoothing and the issue of heteroskedasticity as pointed out by Sims (2001) and Sims and Zha (2002).[6]

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Our results indicate that there is significant time variation in the monetary policy behavior of the European countries that has not been accounted for by the fixed-coefficient approach of CGG. We find that the Bundesbank's response to expected inflation started increasing in 1978 and reached its peak in 1987. There was a slight decrease in the response during the reunification of Germany, which is supported by the historical evidence. We also find that the estimate of the response coefficient was more than one for one for the whole sample period. This is consistent with the image of the Bundesbank as an implicit inflation targeter and its success in containing inflation in the late 1970s and the early 1980s.[7] Our findings suggest that the Bank of England adopted an aggressive anti-inflationary stance in the early 1980s, as its response to inflation became significantly higher than one, and the response to output gap was negative and insignificant. This coincides with the disinflation regime under Thatcher [Batini and Nelson (2005)]. The fixed-coefficient monetary policy rule is not able to capture this time variation. Our results confirm the finding that German monetary policy played a significant role in the formulation of monetary policy by the Bank of France and the Bank of Italy. CGG found that the Bank of France's response to changes in the German interest rate was greater than the corresponding response of the Bank of Italy. Our findings suggest that it masks significant time variation in the responses of these two central banks to changes in the German interest rate. We find that the Bank of Italy's response to the German interest rate was more aggressive than its French counterpart between 1982 and 1987. There was a significant decline in the response to the German interest rate after 1992, and this coincided with the breakdown of the exchange rate management system in Europe. The responses of the French and Italian central banks to changes in the German interest rate during the exchange rate management system suggests that the main objective of the monetary policy of these two central banks was to maintain the exchange rate band with respect to the German mark.

The structure of the paper is as follows. In Section 2, we explain the empirical model presented in this paper. Section 3 documents the data; Section 4 explains the empirical results of this paper; and Section 5 concludes.

2. MODEL SPECIFICATION

In this paper, we estimate a time-varying monetary policy rule that is based on CGG. CGG considered a forward-looking version of the monetary policy reaction function where the nominal short-term interest rate responds to changes in expected inflation and expected output gap. If r_t denotes the nominal short-term interest rate, the postulated monetary policy reaction function is

$$r_t = (1 - \beta_{3,t})\{\beta_{0,t} + \beta_{1,t} E[\pi_{t,k}|\psi_t] + \beta_{2,t} E[g_{t,q}|\psi_t]\} + \beta_{3,t} r_{t-1} + v_t, \quad (1)$$

where $\pi_{t,k}$ denotes the percent change in the price level between periods t and $t+k$ (expressed in annual rates). $g_{t,q}$ is the measure of the average output gap between

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period t and $t+q$, with the output gap being defined as the percent deviation between actual GDP and the corresponding target, which is actually the potential GDP. ψ_t is the information set available at time t. The lagged interest rate is added to account for interest rate smoothing. Equation (1) is different from the target nominal short-term interest rate equation of CGG, as we allow the coefficients to vary with time.

If the observed values of inflation and GDP gap are used in place of the unobserved rational forecasts in equation (1), the model is modified in the following way:

$$r_t = (1-\beta_{3,t})\{\beta_{0,t} + \beta_{1,t}\pi_{t,k} + \beta_{2,t}g_{t,q}\} + \beta_{3,t}r_{t-1} + e_t, \tag{2}$$

where $e_t = -(1-\beta_{3,t})\{\beta_{1,t}(\pi_{t,k} - E[\pi_{t,k}|\psi_t]) + \beta_{2,t}(g_{t,q} - E[g_{t,q}|\psi_t])\} + v_t$, and where $\pi_{t,k}$ and $g_{t,q}$ are the actual future values of inflation and output gap, and $(\pi_{t,k} - E[\pi_{t,k}|\psi_t])$ and $(g_{t,q} - E[g_{t,q}|\psi_t])$ are rational forecast errors, which are assumed to be uncorrelated with the rational forecasts. In equation (2), we can easily see that the explanatory variables $\pi_{t,k}$ and $g_{t,q}$ are correlated with the disturbance term e_t. Because the disturbance term e_t is a linear combination of rational forecast errors, any information available at time $t-1$ will also be uncorrelated with the disturbance term e_t. Following CGG (2001) and Kim and Nelson (2006), we assume $k=1$, $q=1$ for our baseline model. The estimation methodology presented in Kim and Nelson (2006) is not suitable for more than a one quarter–ahead targeting horizon, as multi-period ahead–inflation and output gap will induce a moving average component in the error term. Following Boivin (2006) and Kim and Nelson (2006), we allow the coefficients assumed to follow a random walk:

$$\beta_{i,t} = \beta_{i,t-1} + \varepsilon_{i,t}, \quad \varepsilon_{i,t} \sim \text{i.i.d. } N\left(0, \sigma^2_{\varepsilon,i}\right), \quad i = 0, 1, 2, 3. \tag{3}$$

There are alternative approaches to modeling time variation that includes structural breaks as well as Markov switching in the reaction function coefficients. The usual test of time variation has low power against the alternative; that is, it is difficult to distinguish between different forms of time variation. As in Boivin (2006), we note that structural break models are very special cases of time variation and do not allow for the gradual evolution of monetary policy. Moreover, a time-varying parameter model may also be used as a good approximation to multiple breaks in the reaction function coefficients.[8]

Furthermore, following Sims (1999) and Sims and Zha (2002), we allow for heteroskedasticity in the disturbance term e_t by assuming a GARCH (1,1) process,

$$e_t|\psi_{t-1} \sim N\left(0, \sigma^2_{e,t}\right), \tag{4}$$

$$\sigma^2_{e,t} = \alpha_0 + \alpha_1 e^2_{t-1} + \alpha_2 \sigma^2_{e,t-1}. \tag{5}$$

The maximum-likelihood estimate of the model in equations (2)–(5) via the conventional extended Kalman filter would result in invalid inferences, as the regressors of the model are correlated with the disturbance term. In this paper,

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we adopt the approach proposed by Kim (2006) and Kim and Nelson (2006) for consistent and efficient estimation of the model presented in equation (2).[9] To take care of the endogeneity problem, inflation and output gap are regressed on a set of instruments, and standardized residuals v_{1t}^* and v_{2t}^* are obtained for inflation and output gap. The error terms in the first-stage regression are also heteroskedastic. Kim and Nelson (2006) have shown that an equivalent way of writing the error term in equation (2) is

$$e_t = \rho_1 \sigma_{e,t} v_{1t}^* + \rho_2 \sigma_{e,t} v_{2t}^* + \omega_t,\ \omega_t \sim N\left(0, \left(1 - \rho_1^2 - \rho_2^2\right)\sigma_{e,t}^2\right), \qquad \mathbf{(6)}$$

where ω_t is uncorrelated with v_{1t}^* and v_{2t}^*. The idea is to decompose the error term of the monetary policy rule, equation (2), into two components: $\mathrm{E}[e_t|\psi_t] = (\rho_1 \sigma_{e,t} v_{1t}^* + \rho_2 \sigma_{e,t} v_{2t}^*)$, which is correlated with the explanatory variables, and $\omega_t = e_t - \mathrm{E}[e_t|\psi_t]$, which is uncorrelated with the explanatory variables. If there is no endogeneity in equation (2), the coefficients ρ_1 and ρ_2 should be insignificant. Substituting equation (6) into equation (2), we get

$$r_t = (1 - \beta_{3,t})\{\beta_{0,t} + \beta_{1,t}\pi_{t,1} + \beta_{2,t} g_{t,1}\} + \beta_{3,t} r_{t-1} + \rho_1 \sigma_{e,t} v_{1t}^* + \rho_2 \sigma_{e,t} v_{2t}^* + \omega_t, \qquad \mathbf{(2')}$$

where $\omega_t \sim N(0, (1 - \rho_1^2 - \rho_2^2)\sigma_{e,t}^2)$. The augmentation of the monetary policy rule equation with the standardized bias error correction term takes care of the endogeneity problem in equation (2). Using the standardized forecast error terms as bias correction terms in equation (2′) also takes care of the heteroskedasticity in the first-stage regression, as our estimates in the second step will be inefficient if we do not take heteroskedasticity into account in the first stage regression.

It is also well known that a general Taylor rule such as equation (1) does not capture the behavior of all central banks. In this paper, we assume that the Bundesbank and the Bank of England have followed independent monetary policies; however, there has been consensus in the literature that the Bank of France and the Bank of Italy followed the actions of the Bundesbank during the exchange rate management system. Therefore, we need to augment the monetary policy rules of France and Italy by including the German interest rate. The modified short-term interest rate targets for these two countries become

$$r_t^* = \beta_{0,t} + \beta_{1,t} E[\pi_{t,k}|\psi_t] + \beta_{2,t} E[g_{t,q}|\psi_t] + \beta_{4,t} E[y_{t,p}|\psi_t], \qquad \mathbf{(7)}$$

where $y_{t,p}$ is the German interest rate between periods t and $t + p$. We follow the same two-step estimation procedure to estimate the above equation. The only difference is that the set of instruments is expanded to include lags of y_t in the estimation of this augmented monetary policy rule equation. As in CGG, we choose $p = 0$.

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3. DATA DESCRIPTION

We use quarterly data from the International Financial Statistics (IFS), starting in the first quarter of 1973. The sample runs through the last quarter of 1998 for France, Germany, and Italy. We need to truncate the sample for these countries at the end of 1998, as they joined the European Monetary System (EMS) at the beginning of 1999. For the United Kingdom, the sample extends through the first quarter of 2008. The interest rate is the respective overnight interbank lending rate, which is the call money rate in the IFS database. Inflation rate is calculated using the consumer price index (CPI) of each country, and it is quarterly annualized changes in the CPI. The output gap is the cyclical component, obtained by the Hodrik–Prescott (HP) filter.[10] We decompose the log of real GDP for each country, except Italy, to obtain the output gap. In the case of Italy, real GDP data that go back to 1973 are not available; therefore, we use the index of industrial production (IIP) as a proxy for output. The set of instruments includes one to four lags of output gap, inflation, interest rates, and commodity inflation. Our sample period is longer and more consistent for each country than that of CGG, who use different sample periods for different countries.

In this paper, we assume that Germany and the United Kingdom followed an independent monetary policy for the period under study. This assumption implies that monetary policy formulation by their central banks was not affected by the actions of other countries. We think that this is not a very strong assumption, as other papers in the literature have also made this assumption in estimating the monetary policy reaction functions of these two countries.[11] All the countries under study were members of the Exchange Rate Mechanism (ERM) of the EMS, although for varying time periods. The Bundesbank is regarded as the leader in the ERM, setting its monetary policy independently, and not responding to movements in its exchange rate. The United Kingdom was a member of the ERM briefly between October 1990 and September 1992. It has been argued that the Bank of France and the Bank of Italy were constrained by the EMS commitments. Movements in the German interest rate played a significant role in the monetary policy formulation of these two central banks. Therefore, we also estimate the monetary policy reaction functions augmented with the German interest rates for the Bank of France and the Bank of Italy.

4. EMPIRICAL RESULTS

The smoothed time-varying response coefficients for the Bundesbank are shown in Figure 1. The estimated hyperparameters are shown in Table 1. The coefficients on the bias error correction terms, ρ_1 and ρ_2, are negative and jointly significant at the 1% level. Thus, ignoring endogeneity in the estimation of the reaction function will yield biased estimates. The inflation response coefficient conforms to the Bundesbank's image as an anti-inflation hawk. The Bundesbank increased the short-term interest rate by more than 100 basis points in response to a 1%

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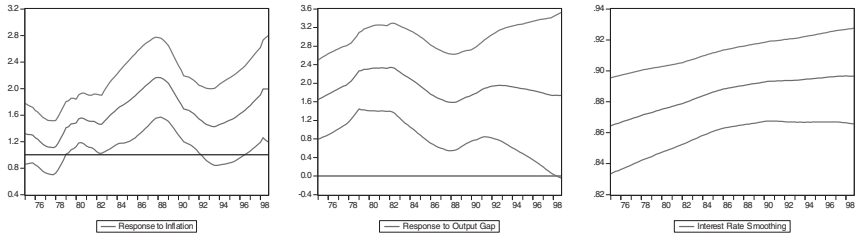


FIGURE 1. Germany (baseline model).

increase in expected inflation for the whole sample period. Our results show that the Bundesbank's response to expected inflation started increasing before 1978, and it became significantly higher than one after 1979. This is consistent with the evidence presented in Cecchetti et al. (2007), who find that the Bundesbank was able to stabilize inflation much before the United States and the United Kingdom. The response to expected inflation declined marginally in the early 1990s, and was not significantly different from one for some time period. This was the period of German reunification, and inflation was briefly above 5%. The response to the output gap was significant for the whole sample period. Our results show that the response to the output gap peaked in the last quarter of 1978, and it declined at a very slow pace till 1988. The decline in the output gap coefficient and the increase in response to the movements in expected inflation in the early 1980s were instrumental in inflation stabilization occurring in Germany before

TABLE 1. Estimation of hyperparameters for the baseline monetary policy rule

Parameter	Germany	U.K.	France	Italy
α_0	0.123 (0.045)	0.095 (0.030)	0.095 (0.045)	0.423 (0118)
α_1	0.040 (0.072)	0.051 (0.082)	0.011 (0.060)	0.011 (0.115)
α_2	0.902 (0.123)	0.865 (0.070)	0.842 (0.163)	0.930 (0.122)
$\sigma_{\varepsilon,0}$	0.001 (0.003)	0.002 (0.005)	0.969 (0.232)	0.001 (0.002)
$\sigma_{\varepsilon,1}$	0.140 (0.051)	0.242 (0.068)	0.229 (0.079)	0.107 (0.038)
$\sigma_{\varepsilon,2}$	0.193 (0.112)	0.864 (0.122)	0.091 (0.212)	0.326 (0.206)
$\sigma_{\varepsilon,3}$	0.024 (0.021)	0.773 (0.146)	0.205 (0.072)	0.251 (0.087)
ρ_1	−0.276 (0.088)	−0.127 (0.075)	0.010 (0.106)	−0.392(0.103)
ρ_2	−0.140 (0.102)	0.314 (0.073)	0.019 (0.133)	−0.259 (0.095)

Note: Standard errors are in parentheses. The assumed model is

$$r_t = (1-\beta_{3,t})\{\beta_{0,t} + \beta_{1,t}\pi_{t,1} + \beta_{2,t}g_{t,1}\} + \beta_{3,t}r_{t-1} + e_t, e_t \sim N\left(0, \sigma^2_{e,t}\right),$$

$$e_t = \rho_1\sigma_{e,t}v^*_{1t} + \rho_2\sigma_{e,t}v^*_{2t} + \omega_t,\ \omega_t \sim N\left(0, \left(1-\rho_1^2-\rho_2^2\right)\sigma^2_{e,t}\right),$$

$$\sigma^2_{e,t} = \alpha_0 + \alpha_1 e^2_{t-1} + \alpha_2\sigma^2_{e,t-1},$$

$$\beta_{i,t} = \beta_{i,t-1} + \varepsilon_{i,t},\ \varepsilon_{i,t} \sim \text{i.i.d. } N\left(0, \sigma^2_{\varepsilon,t}\right),\quad i = 0, 1, 2, 3.$$

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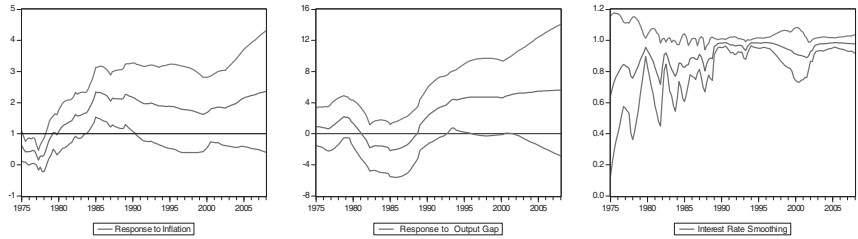


FIGURE 2. United Kingdom (baseline model).

the United States and the United Kingdom. There was a slight increase in the response to output gap during the German reunification period, which is supported by historical evidence. The evidence presented here clearly finds support for time variation in the response of the Bundesbank to changes in expected inflation and output gap over time.

Figure 2 shows the estimated response coefficients for the Bank of England in a baseline model. The estimated hyperparameters are shown in Table 1. As shown in the figure, the short-term interest rate increased by less than 100 basis points in response to a 1% increase in expected inflation before 1979. This suggests that the U.K. monetary policy failed to provide a nominal anchor in the pre-1979 period. The point estimate of the inflation response coefficient was higher than one for the post-1979 sample period; however, it became significantly higher than one only after 1983.[12] This is in contrast to the German case, where the Bundesbank's response to inflation was significantly higher than one starting in 1979 itself. The response to expected inflation in the United Kingdom peaked around 1990, and the point estimate declined marginally until 1998, although it was consistently higher than one. There was a widening of the confidence band around the inflation response coefficient in the 1990s. This is in contrast to the fixed coefficient approach, where CGG find that the average response of the Bank of England to a 1% increase in expected inflation was 0.98% for the sample period 1979–1990. The point estimate of the output gap response coefficient shows the anti-inflation stance of the Bank of England in the early 1980s, as it was negative and insignificant between 1981 and 1987. The Bank of England started putting more weight on output stabilization once inflation was stabilized. The response to output gap became significant after 1987.

Why did we see a fundamental change in the way British monetary policy responded to changes in inflation and output gap after 1979? One of the explanations may be the election of the Margaret Thatcher government in 1979, and its emphasis on inflation stabilization. The Bank of England undertook monetary policy by adopting a medium-term financial strategy in 1979, although it was abandoned in October 1985. The shift in the monetary policy in the United Kingdom is similar to what Kim and Nelson (2006), and Kim et al. (2007) find for the United States using ex post and real-time data.

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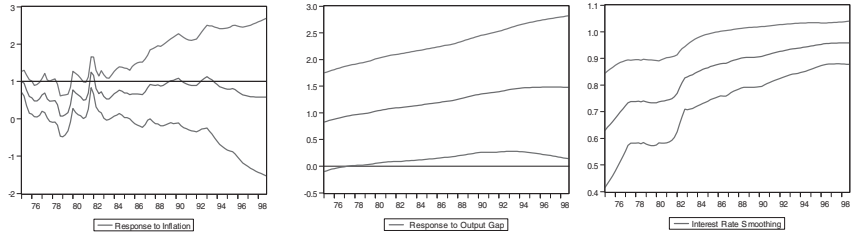


FIGURE 3. France (baseline model).

The baseline smoothed estimates of coefficients on inflation and the output gap in France are shown in Figure 3. Table 1 shows the estimated hyperparameters for the Bank of France's baseline reaction function.[13] Our estimates show that for the full sample under our study, the French short-term interest rate moved less than one for one in response to changes in expected inflation. The coefficient on the output gap was insignificant in the early part of the sample. The output gap response coefficient became significant after 1979.

According to the conventional view, the Bank of France and the Bank of Italy followed the moves of the Bundesbank very closely. To take this issue into account, we add the German interest rate to the baseline monetary policy reaction function of the Bank of France (see Figure 4). The estimated inflation response coefficient is less than one in this case. This does not imply that the effective weight on inflation declined after the inclusion of the German rate in the monetary policy reaction function. If the estimated coefficient on inflation is 0.5 in a model where the coefficient on the German interest rate is 0.6, this is equivalent to an effective weight of 1.25 on inflation in the baseline model ($1.25 \times 0.4 = 0.5$).[14] If the German interest rate is added to the baseline specification, we find that the Bundesbank's interest rate–setting behavior heavily influenced the decisions of the Bank of France until 1992. In fact, a one–percentage point rise in the German rate led to more than a 100–basis point increase in the French interest rate for some time periods under study. The reaction to the German interest rate declined over time and became insignificant for the major part of the post-1992 period. For the initial sample period, our result is consistent with CGG, who find a 114–basis point rise in the French interest rate in response to a one percentage increase in the German interest rate. However, our results find substantial time variation in

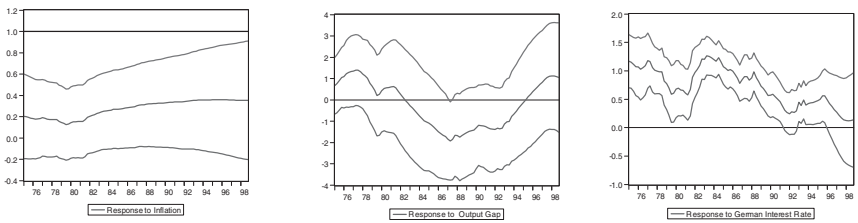


FIGURE 4. France (augmented model).

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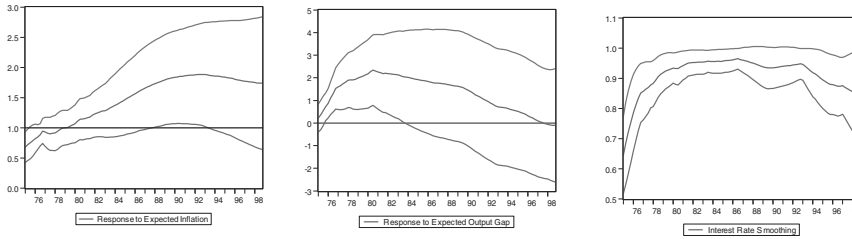


FIGURE 5. Italy (baseline model).

response to the German rate and we find that the degree of follow-up to the German rate declined substantially over time. This is an interesting result, as it shows that the Bank of France's main objective for monetary policy during the pre-1992 period was to follow the actions of the Bundesbank very closely. This was also the time period when France was under the Exchange Rate Management System (EMS). The big decline in the response to the German interest rate coincides with the breakup of the EMS regime in Europe.

The estimates of the baseline reaction function for Italy (see Figure 5) indicate that the Bank of Italy's response to expected inflation was more than one for one after 1979, although not significantly different from one for most of the sample period. The response to the output gap was also positive during the whole time period of our study. The estimated response coefficients indicate that the Bank of Italy followed Taylor's prescription in responding to fluctuations in inflation and output gap. But these coefficients will give us a misleading answer, if the baseline model is misspecified. This is especially important in the context of Italy, because it did not experience the kind of macroeconomic stability the United States and Germany did after 1982.

It is well known that the Bank of Italy followed the actions of the Bundesbank very closely during our sample period. We add the German interest rate to the baseline monetary policy rule equation to investigate the importance of German monetary policy in the monetary policy decision making of the Bank of Italy. As shown in Figure 6, the Bank of Italy's monetary policy closely followed the monetary policymaking of its German counterpart. Until 1989, the Bank of Italy's

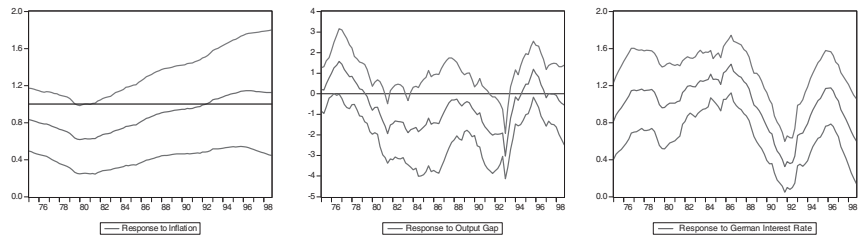


FIGURE 6. Italy (augmented model).

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response to the German interest rate was not significantly different from one. This response coefficient started declining, and reached its minimum in 1992. This is an interesting result that matches the historical narrative of monetary policy in Italy, as it was in 1992 when Italy left the ERM. We also find that there was an increase in the response to the movements in the German interest rate, and it peaked in the last quarter of 1996. This coincided with the reentry of Italy into the ERM in November 1996. In the case of France, the response to the German rate was insignificant for the most part after 1992, even though France stayed in the ERM. This may have been caused by a substantial increase in the exchange rate bandwidth with respect to the German mark from $\pm 2.5\%$ to $\pm 15\%$. In the case of Italy, a severe recession in the immediate aftermath of the collapse of the ERM led to a much stronger response to domestic macroeconomic conditions. Once the macroeconomic conditions stabilized,[15] Italy rejoined the ERM in November 1996 and the response to the German interest rate became significant.

Our results suggest that a fixed-coefficient model is not adequate to capture the gradual evolution of monetary policy in these four European countries. However, we have not undertaken a direct statistical comparison between the TVP model and the constant-coefficient model so far. To test the significance of time variation, we perform a likelihood ratio (LR) test. Table 2 shows the LR statistic for the null of constant coefficients for all four countries under our study. Under the null, we restrict the variances of the time-varying coefficients to be equal to zero. The critical value for the LR test under a standard distribution assumption at a 5% significance level is 9.49 for Germany and the United Kingdom, and 11.08 for France and Italy.[16] However, we know that the presence of a nuisance parameter under the null leads to overrejection of the null hypothesis [Hansen (1992); Andrews and Ploberger (1994)], and this leads to critical values that are too low under the standard distribution assumption. Because the LR statistic for all countries is significantly higher than the standard critical value, there is strong evidence against the null of no time variation.

Our baseline specification is based on CGG's model (2000), where the targeting horizon is the one–quarter ahead inflation and output gap. To examine whether our results are sensitive to the choice of targeting horizons, we estimate the monetary policy reaction function with two alternative targeting horizons, $k = 1$, $q = 0$, represented by [1, 0] and $k = 0$, $q = 0$, represented by [0, 0]. We consider

TABLE 2. Likelihood ratio (LR) statistic for test of time variation

Country	LR stat.
Germany	29.82
U.K.	36.34
France	63.54
Italy	65.38

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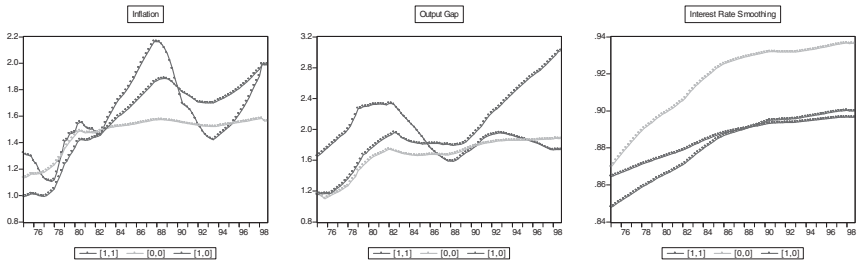


FIGURE 7. TVP estimates for different policy horizons (Germany).

one–quarter ahead inflation and current-quarter output gap because the monetary policy transmission lag is longer for prices than for real activity.[17] The two-step bias correction method is used to estimate the $k = 1$, $q = 0$ case, whereas the standard Kalman filter can be applied to estimate the reaction function for $k = 0$, $q = 0$ (current quarter inflation and output gap) targeting horizon.

Figures 7–10 compare the estimated baseline coefficients with alternative targeting horizons. Even though the broad pattern of time variation is similar, we find that response to current quarter inflation is smaller than one–quarter ahead inflation for all countries.[18] The results suggest that the monetary policy response coefficient for baseline specification [1, 1] is closer to the one–quarter ahead inflation and current output gap targeting horizon than the current-quarter targeting horizon for

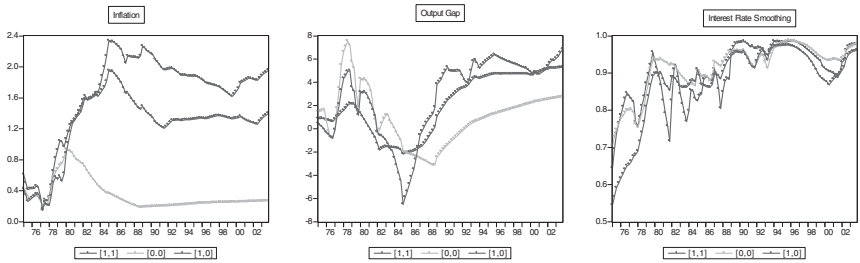


FIGURE 8. TVP estimates for different policy horizons (United Kingdom).

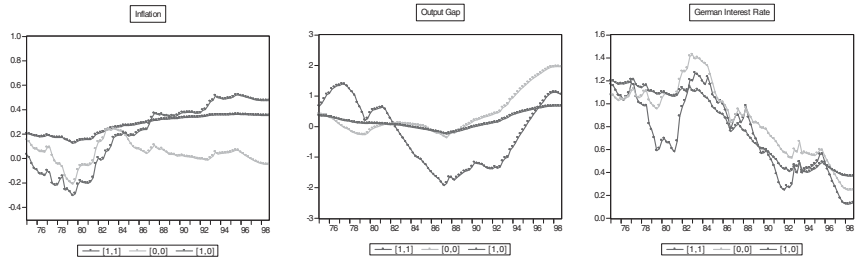


FIGURE 9. TVP estimates for different policy horizons (France).

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FIGURE 10. TVP estimates for different policy horizons (Italy).

both the inflation and the output gap. The results presented here clearly show the usefulness of estimating a forward-looking monetary policy rule. The response to movements in the German interest rate in the cases of France and Italy follows a similar pattern for different targeting horizons, though there is less time variation when the targeting horizon is [0, 0].

It has also been suggested that the exchange rate may have played a significant role in the formulation of monetary policy in these European countries. If we augment the baseline model with exchange rates, we do not find a significant response of the interest rate to changes in the exchange rate. Therefore, we do not report the results for the forward-looking monetary policy rule with exchange rate. The nonsignificance of exchange rate, however, is not surprising. As pointed out by Clarida et al. (2001) and Engel and West (2006) the future inflation and output gap already contains information about the exchange rate, and hence in a regression of the interest rate on inflation, output gap, and exchange rate, it is not uncommon to find an insignificant response to exchange rates.

Our estimation results show that the confidence band around the estimated inflation and output gap response coefficients widened in the later part of the sample for all countries. This does not imply that the central banks started paying less attention to the movements in inflation and output gap. As the volatility of inflation and real GDP declined substantially in the 1990s, the short-term interest rate carries less information about inflation and output gap movements. For example, in a regression $y = \beta x + \varepsilon$, $\text{Var}(\widehat{\beta}) = \sigma^2 / \Sigma (x - \overline{x})^2$, if the volatility of x declines, $\text{Var}(\widehat{\beta})$ increases. In an extreme case, when the central bank controls inflation perfectly, the resulting low variation in inflation may lead to insignificant estimates of the policy response to expected inflation.[19]

5. CONCLUSIONS

In this paper, we examine the conduct of the monetary policy in France, Germany, Italy, and the United Kingdom by estimating a time-varying forward-looking monetary policy rule based on Clarida et al. (1998). Following Boivin (2006) and Cogley and Sargent (2001), we note that a fixed-coefficient approach to estimating

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a monetary policy rule is not adequate to understand the conduct of monetary policy in these countries. To estimate the forward-looking monetary policy rule, we utilize the framework developed by Kim (2006) and Kim and Nelson (2006) to deal with the issue of endogeneity in a time-varying parameter model, as the conventional extended Kalman filter would result in invalid inferences. Our results show that the monetary policy behavior of the countries under study has shown significant time variation, and the conventional fixed-coefficient approach is not adequate to explain the gradual evolution of monetary policy in these countries. Our results confirm the conventional wisdom about the Bundesbank as an anti-inflation hawk, as its response to expected inflation for the 1979–1998 time period was more than one for one, although we find that there was a decline in this response during the German reunification period. Our results suggest that the Bank of England adopted an anti-inflationary stance in the early 1980s, as the response to inflation was significantly higher than one, and the response to the output gap was negative and insignificant during this time period. German monetary policy played a significant role in the formulation of monetary policy of the Bank of France and the Bank of Italy. In contrast to the conventional wisdom, our findings suggest that the Bank of France's and the Bank of Italy's responses to the German interest rate declined significantly after 1992, and this coincided with the breakdown of the Exchange Rate Management System in Europe.

NOTES

1. Two sets of data are used in the estimation of the Taylor rule: ex post and real-time. The ex post data method uses the actual future data and the real-time approach uses the data set that was available to the policymakers in real-time. This paper takes the ex post data approach. Orphanides (2001, 2002) and Boivin (2006) adopt the real-time approach. Other notable papers in the literature include Judd and Rudebusch (1998), Clarida et al. (2000), and Kozicki and Tinsley (2007).

2. See also Cukierman (1999) and Ruge-Murcia (2003).

3. Ex post vs. real-time.

4. This problem does not arise in the TVP framework of Boivin (2006), as he uses Greenbook forecast data. Given the assumption that Greenbook forecasts are prepared under the assumption of constant interest rates, these forecasts are uncorrelated with the error term, and therefore a standard Kalman filter can be applied.

5. Partouche (2007) also estimates a time-varying monetary policy rule for the Federal Reserve in a GMM framework.

6. Sims and Zha (2002) argue that a specification for a monetary policy rule ignoring heteroskedasticity in the disturbance term would result in misleading estimates of the time-varying reaction coefficients.

7. See Cecchetti et al. (2007).

8. Stock and Watson (2002) and Boivin (2006) discuss the merits of the TVP model over other forms of structural break.

9. The derivation of a Heckman-type two-step procedure to take into account the endogeneity problem has been adopted from Kim and Nelson (2006).

10. We also perform the analysis using output gap data from quadratic detrending, and the results are qualitatively similar. The results are available upon request.

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11. For example, Assenmacher-Wesche (2006).
12. This is consistent with the findings of Assenmacher-Wesche (2006).
13. The coefficients on the bias error correction terms are not significant for the baseline model; however, they are significant for the augmented model. The results for the augmented model are not presented in this paper, but they are available upon request.
14. This is based on CGG (1998).
15. Wyplosz (1999); Eichengreen and Wyplosz (1993).
16. This test is performed for augmented models of France and Italy.
17. Orphanides (2001) uses the same targeting horizon.
18. Boivin (2006) also finds similar results for the United States.
19. See Kishor and Kochin (2007) for details.

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