

Monetary Policy in the EURO-Area: Was It Too Tight in the 1990s?

Willi Semmler* Alfred Greiner† Wenlang Zhang‡

Abstract

This paper studies the monetary policy in the EURO-area. An important feature of our paper is that we assume that both private behavior as well as the Central Bank's policy responses are time varying. We first investigate empirically why and how the monetary policy of the EURO-area has changed in the past decades. Another important question concerning the EURO-area is whether the monetary policy has been too tight in the 1990s and in turn caused the high unemployment rates, in contrast to the economic prosperity and low unemployment of the U.S.. Therefore the authors explore what would have happened to the EURO-economy if the Central Banks had followed either the fixed or time-varying monetary policy of the U.S.. The paper does find that the European central banks and then later the ECB overreacted to past inflation pressures. We also consider the relationship between the monetary policy and asset prices in the EURO-area.

JEL: E17

1 Introduction

In the profession it is increasingly recognized that formal modeling of monetary policy faces great challenges because of (1) uncertainty of what the model should look like that monetary authorities presume private agents are following (model uncertainty), (2) uncertainty about the actual situation of the economy (data uncertainty), (3) uncertainty about the size of shocks (shock uncertainty), and (4) uncertainty concerning the effects of monetary policy actions (uncertainty about intended and unintended consequences of actions). Thus, it is of no surprise that neither private agents' reactions to monetary policy actions nor monetary policy behaviors have been stable over time. Assumed models, underlying private behavior reaction coefficients of basic economic relations and monetary policy behavior change over time. Moreover, what has been called

*Center for Empirical Macroeconomics, Bielefeld, and New School University, New York.

†Center for Empirical Macroeconomics, Bielefeld University, Germany.

‡Center for Empirical Macroeconomics, Bielefeld University, Germany.

the “NAIRU” has changed over time, too. The view that private reactions to monetary policy have changed over time is frequently associated with the Lucas critique. Lucas has maintained that response coefficients of private agents to monetary policy will not be constant over time. Researchers have characterized monetary regimes where economic and monetary relationships as well as monetary policy are relatively stable only for some time period and subsequently followed by shifts.

An important shift in monetary regime, as far as the U.S. is concerned, has been observed at the beginning of the 1980s when Paul Volcker was appointed chairman of the Fed. This shift was seen as a shift from passive monetary policy with strong reaction to employment but passive reaction to inflation, to an active anti-inflation policy, namely monetary policy strongly reacting to inflation. Similar shifts have also been observed for the EURO-area at that time period. On the other hand, in the U.S. there seems to be a shift back to lower interest rates since the beginning of the 1990s, giving more weight to the output gap and less to the inflation gap. The shift in the latter period does not seem to have been followed by the EURO-area economy.

An interesting question is thus whether the monetary policy of the EURO-area economy has been too tight in the 1990s and therefore caused high unemployment rates and low economic growth rates, in contrast to the prosperity and low unemployment rate of the U.S.. Therefore the question can be asked, what would have happened if the EURO-area economy had followed the U.S. monetary policy.

In order to take account of the Lucas critique the authors estimate the time varying Phillips curve in section 2. This concerns the time varying reaction of the private sector to the unemployment gap as well as the time variation of what has been called the natural rate. In section 3 a monetary policy model is set up with a central bank objective function to derive the optimal interest rate reaction (Taylor rule), and estimations for the EURO-area economy as well as for the U.S. time varying response coefficients for the Taylor rule are undertaken. The paper estimates the time varying parameters by way of the Kalman filter. Section 4 employs the U.S. time varying and fixed response coefficients of the Taylor rule and simulates the time paths of the EURO-area interest rate, inflation rate and output gap. In section 5, we explore the relationship between the monetary policy and asset price volatility. Section 6 concludes the paper.

2 Time Varying Phillips Curve

2.1 Time Varying Reaction to the Unemployment Gap

This subsection estimates the traditional Phillips-Curve¹

$$\pi_t = \alpha_0 + \sum_{i=1}^n \alpha_i \pi_{t-i} + \alpha_{ut}(U_t - U_t^N) + \xi_t, \quad (1)$$

$$\alpha_{ut} = \alpha_{ut-1} + \eta_t, \quad (2)$$

where π_t is the inflation rate, U_t the unemployment rate and U_t^N the so-called NAIRU. ξ_t and η_t are both i.i.d. with mean zero and variance σ_ξ^2 and σ_η^2 respectively. The number of lags depends on the t-statistics of the corresponding coefficients, namely, if the t-statistics are insignificant this lag of inflation rate will be excluded. Equation (2) assumes that α_{ut} is time dependent and takes a random walk path. In order to estimate the time varying path of α_{ut} , we take advantage of maximum likelihood estimation by way of the Kalman filter. The countries of our study include Germany, France, the U.K., Italy and the U.S.. Quarterly data are used. The data for all countries are taken from International Statistics Yearbook 2000 (online). T-statistics are included in parentheses.²

The inflation rate of Germany is measured by changes in the Consumer Price Index (CPI). The NAIRU is assumed to be fixed at 6 percent. This is undoubtedly a simplification, since the NAIRU may also change over time.³ From the data for 1963:4-98:4 we obtain the following results:

$$\pi_t = 0.005 + 1.047\pi_{t-1} - 0.181\pi_{t-2} + \alpha_{ut}(U_t - U_t^N).$$

(1.495) (9.922) (2.268)

The path of α_{ut} is presented in Figure 1A.

The inflation rate of France is measured by the log difference of the GDP deflator. The NAIRU is also assumed to be 6 percent. The data cover 1969:1-99:4 and the estimation reads

$$\pi_t = 0.008 + 0.901\pi_{t-1} - 0.003\pi_{t-2} + \alpha_{ut}(U_t - U_t^N).$$

(0.566) (6.070) (0.045)

The path of α_{ut} is presented in Figure 1B.

The inflation rate of the U.K. is measured by changes in the CPI. The NAIRU is assumed to be 6 percent. The data cover 1964:1-99:4, the result is

$$\pi_t = 0.007 + 1.384\pi_{t-1} - 0.491\pi_{t-2} + \alpha_{ut}(U_t - U_t^N).$$

(2.403) (15.845) (6.695)

¹In our estimation of the Phillips curve we take a backward-looking approach as justified in Rudebusch and Svensson (1999).

²The use of the Kalman filter for our propose is discussed in Semmler, Greiner and Zhang (2001). For a more technical description, see Hamilton (1994).

³In this subsection, we assume that the NAIRU is fixed for all countries and take a value which is close to the average unemployment rate for the period covered. In the next subsection, we estimate the time varying NAIRU.

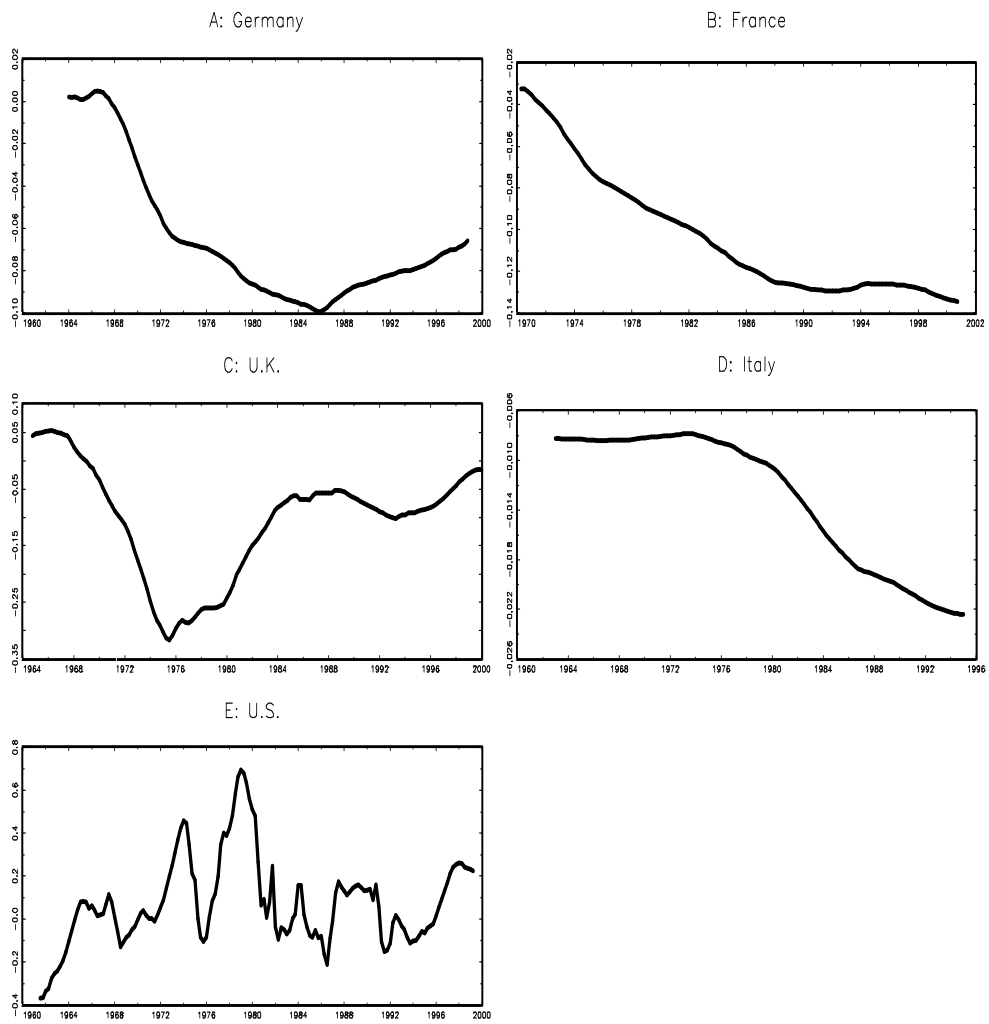


Figure 1: α_{ut} for Germany, France, the U.K., Italy and the U.S.

The path of α_{ut} is presented in Figure 1C.

The inflation rate of Italy is measured by changes in the CPI and the NAIRU is taken as 5 percent. For the data 1962-99, the changes of α_{ut} are insignificant, but for the period 1962-94 the changes are significant enough, therefore the estimation is undertaken for 1962:3-94:3 with the following result

$$\pi_t = 0.004 + 1.409 \pi_{t-1} - 0.448 \pi_{t-2} + \alpha_{ut}(U_t - U_t^N).$$

(0.887) (14.111) (2.870)

The path of α_{ut} is presented in Figure 1D.

Besides the estimations for the four European countries, we also undertake estimation for the U.S.. The inflation rate of the U.S. is measured by changes in the CPI. The NAIRU is taken to be 5 percent. The data cover 1961:1-99:4 and the result is

$$\pi_t = 0.004 + 1.198 \pi_{t-1} - 0.298 \pi_{t-2} + 0.203 \pi_{t-3} - 0.202 \pi_{t-4} + \alpha_{ut}(U_t - U_t^N).$$

(2.665) (12.242) (2.119) (1.589) (2.275)

The path of α_{ut} is presented in Figure 1E. In Figure 1E, we find that for many years α_{ut} is positive, which is inconsistent with the traditional view that there is a negative relationship between inflation rate and unemployment gap. One reason seems to be related to the value of the NAIRU, which is assumed to be fixed at 5 percent here. In fact, higher values of the NAIRU have also been tried and it seems that the higher the NAIRU, the smaller the value of α_{ut} .

From the evidence above, we find that the α_{ut} in eq.(1) does experience changes. For the three EU countries Germany, France and Italy, we find that the changes of α_{ut} are to some extent similar. α_{ut} has been decreasing persistently since the 1960s for France and Italy. In the case of Germany, it has been increasing slowly from the middle of the 1980s on. As for the U.K., the change of α_{ut} is relatively different from those of the other three countries. It decreased very fast in the 1960s and started to increase in 1975. In order to analyze the causes of the differences of the evolution of α_{ut} , we present the inflation and unemployment rates of the four EU countries from 1970 to 1999 in Figure 2 and 3 respectively. It is obvious that the changes of inflation rates of the four countries are similar. π_t attained its highest point around 1975, decreased to a low value in about 4 years, increased to another peak at the end of the 1970s and then went down until 1987, after which it evolved smoothly and stayed below 10 percent. The evolution of inflation rate does not seem to be responsible for the different paths of α_{ut} of the four countries. The evolution of the unemployment rates in Figure 3, however, may partly explain why the change of α_{ut} in the U.K. is somewhat different from those of the other three countries. Before 1986 the unemployment rates of the 4 countries have been going up on the whole, while after 1986, there exist some differences. The evolution of U_t in the U.K. is not very consistent with those of the other three countries. After 1992, U_t of the U.K. decreased relatively fast from about 10 percent to 4 percent, while those of the other three states remained relatively high during the whole 1990s and did not begin to go down until 1998.

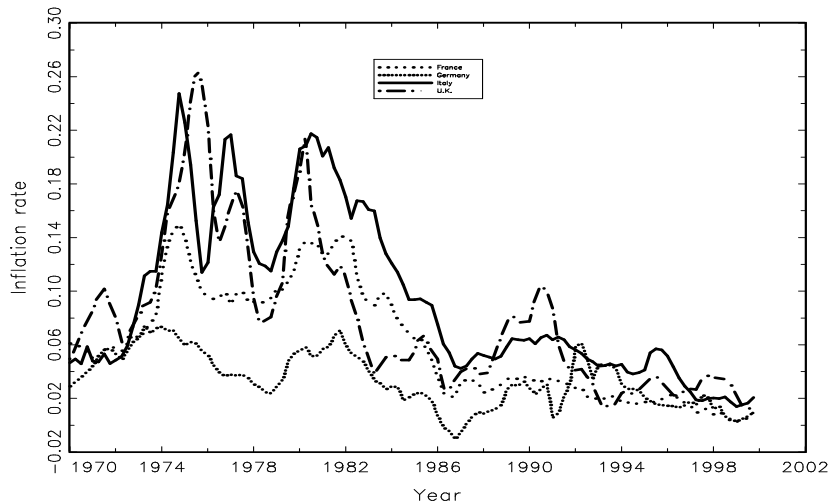


Figure 2: Inflation Rates of Germany, France, Italy and the U.K.

2.2 Time Varying NAIRU with Supply Shocks

In the previous subsection, we have assumed that the so-called NAIRU is fixed. Next we are interested in the question whether the NAIRU has changed over time. A comprehensive description of how the NAIRU can be estimated can be found in Staiger, Stock and Watson(1996). This section reports some estimates of the time varying NAIRU for the five countries: Germany, France, the U.K., Italy and the U.S. with a model adopted from Gordon (1997):

$$\pi_t = a(L)\pi_{t-1} + b(L)(U_t - U_t^N) + c(L)z_t + e_t, \quad (3)$$

$$U_t^N = U_{t-1}^N + \eta_t, \quad (4)$$

where π_t is the inflation rate, U_t the actual unemployment rate and U_t^N the NAIRU, which follows a random walk path indicated by equation (4). z_t is a vector of supply shock variables. L is a polynomial in the lag operator. e_t is a serially uncorrelated error term and η_t satisfies the Gaussian distribution, with mean zero and variance σ_η^2 . Obviously, the variance of η_t plays an important role in the estimation. If it is zero, then the NAIRU is constant, and if it is positive, the NAIRU experiences changes. If no constraints are imposed on σ_η^2 , the NAIRU will jump up and down and soak up all the residual variation in the inflation variation. This is a standard “stochastic time varying parameter regression model” that can be estimated using maximum likelihood methods with

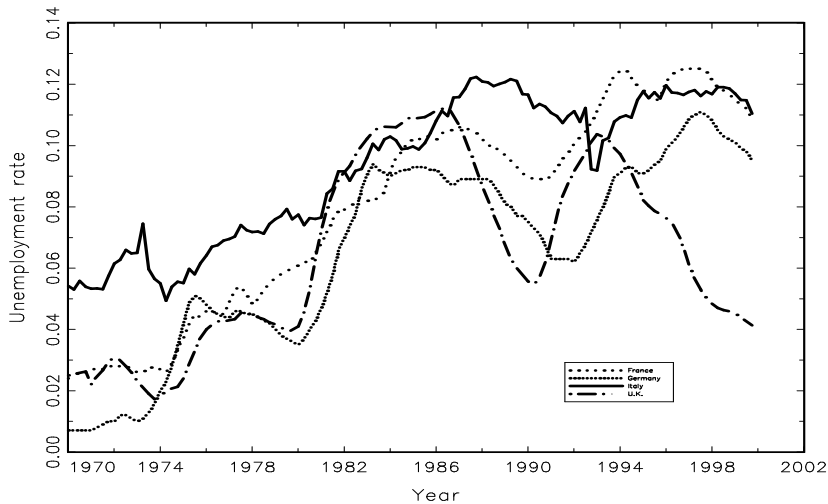


Figure 3: Unemployment Rates of Germany, France, Italy and the U.K.

the help of the Kalman filter, as already analyzed in the previous subsection.

Gordon (1997) includes a z_t to proxy supply shocks such as changes of relative prices of imports and the change in the relative price of food and energy. If no supply shocks are taken into account, the NAIRU is referred to as “estimated NAIRU without supply shocks”. Though there are no fixed rules on what variables should be included as supply shocks, it seems more reasonable to take supply shocks into account than not, since there are undoubtedly other variables than the unemployment rate affecting inflation. In this subsection, the supply shocks considered include mainly price changes of import (im_t), food ($food_t$), and fuel, electricity and water ($fuel_t$). As for which variables should be adopted as supply shocks for the individual countries, we undertake an OLS regression for equation (3) before we start the time varying estimation, assuming that NAIRU is constant. In most cases we exclude the variables whose t-statistics are insignificant. The data source is the same as in the previous subsection.

As mentioned above, the standard deviation of η_t plays a crucial role. Gordon (1997) assumed it to be 0.2 percent for the U.S. for the period 1955-96. There is little theoretical background on how large η_t should be, but since the NAIRU is usually supposed to be relatively smooth, we constrain the change of the NAIRU within 4 percent, which is also consistent with Gordon (1997). Therefore we assume different values of η_t for different countries, depending on how large we expect the change of NAIRU to be.

For Germany, the variance of η_t is assumed to be 7.5×10^{-6} , and the price changes of food, import, and fuel, electricity and water are included as supply shocks. The estimates are presented below, with t-statistics in parentheses,

$$\begin{aligned} \pi_t = & 0.004 + 1.052 \pi_{t-1} - 0.256 \pi_{t-2} + 0.008 \pi_{t-3} + 0.013 fuel_{t-1} \\ & (0.056) \quad (10.982) \quad (3.009) \quad (0.153) \quad (1.105) \\ & + 0.061 food_{t-1} + 0.006 im_{t-1} - 0.042 (U_t - U_t^N) + e_t, \\ & (1.994) \quad (0.627) \quad (0.993) \end{aligned}$$

where $fuel_t$ is the price change of fuel, electricity and water, $food_t$ the price change of food and im_t the price change of imports. The estimate of the standard deviation of e_t is 0.006, with t-statistic being 8.506.

Since the unemployment rates of the four EU countries are presented in Figure 3, we present only the time varying NAIRU here. The estimated time varying NAIRU of Germany is presented in Figure 4A.

For France, only one lag of the inflation rate is used in the regression, since the coefficient of the unemployment rate gap tends to zero when more lags of inflation are included. The price changes of food and intermediate goods are included as supply shocks. Three lags of intermediate goods price changes are included to smooth the estimated NAIRU. The result reads as

$$\begin{aligned} \pi_t = & 0.004 + 0.989 \pi_{t-1} - 0.085 food_{t-1} + 0.132 in_t \\ & (0.312) \quad (27.967) \quad (2.058) \quad (2.876) \\ & - 0.090 in_{t-1} + 0.029 in_{t-2} - 0.054 (U_t - U_t^N) + e_t, \\ & (1.313) \quad (0.772) \quad (1.889) \end{aligned}$$

where in_t denotes the price change of intermediate goods. The estimate of the standard deviation of e_t is 0.005, with t-statistic being 8.808, and the variance of η_t is predetermined as 1.3×10^{-5} . The estimated NAIRU of France is presented in Figure 4B.

Because of the same reason as for France, one lag of inflation is included in the regression for the U.K.. The estimation reads

$$\begin{aligned} \pi_t = & 0.005 + 0.818 \pi_{t-1} + 0.130 food_{t-1} \\ & (0.066) \quad (11.063) \quad (2.768) \\ & + 0.017 fuel_{t-1} - 0.072 (U_t - U_t^N) + e_t \\ & (0.691) \quad (0.300) \end{aligned}$$

The estimate of the standard deviation of e_t is 0.013, with t-statistic being 8.764, and the variance of η_t is predetermined as 1.4×10^{-5} . The estimated NAIRU of the U.K. is presented in Figure 4C.

For Italy it seems difficult to get a smooth estimate for the NAIRU if we include price changes of food, fuel, electricity and water and imports as supply shocks, the main reason seems to be that the inflation rate experienced drastic changes and therefore exerts much influence on the estimates of the NAIRU. Therefore, we try to smooth the estimate of the NAIRU by including the current short-term interest rate (r_t) into the regression, which makes the estimated

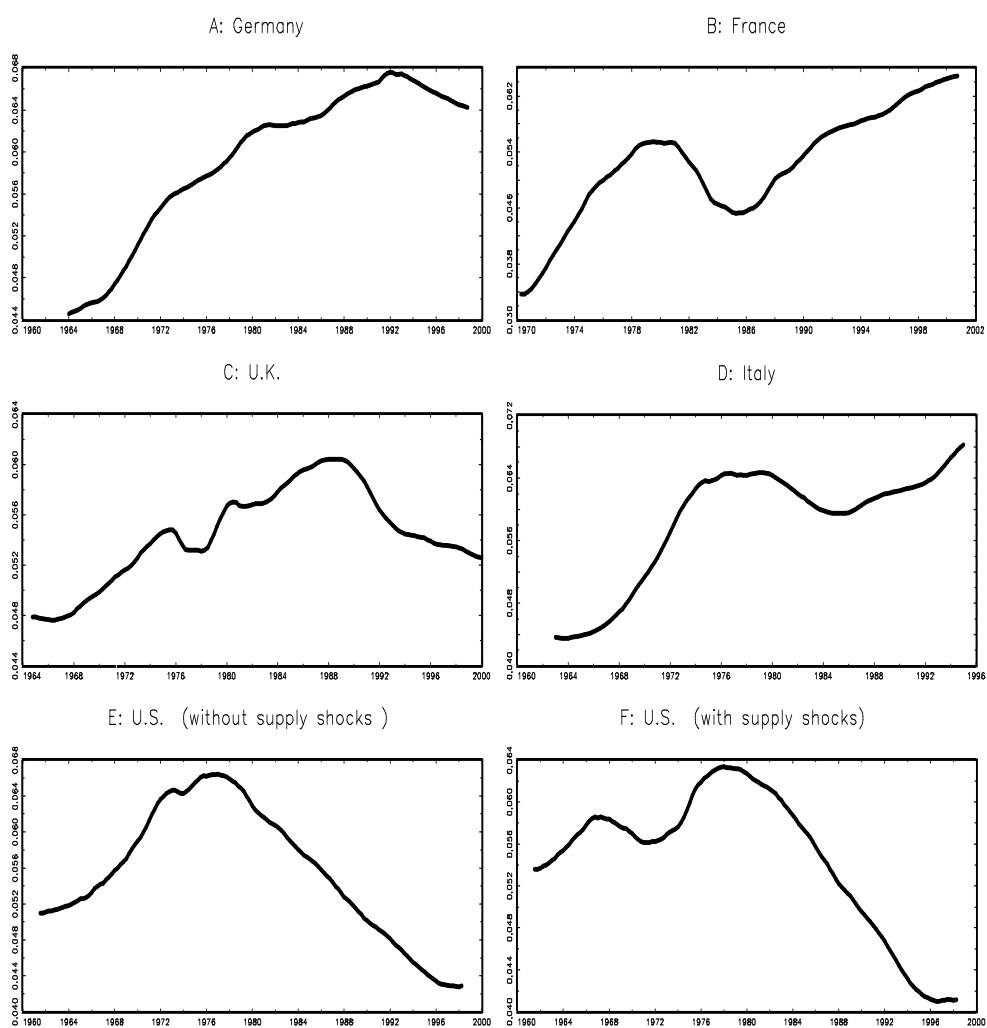


Figure 4: Time varying NAIRU of Germany, France, the U.K., Italy and the U.S.

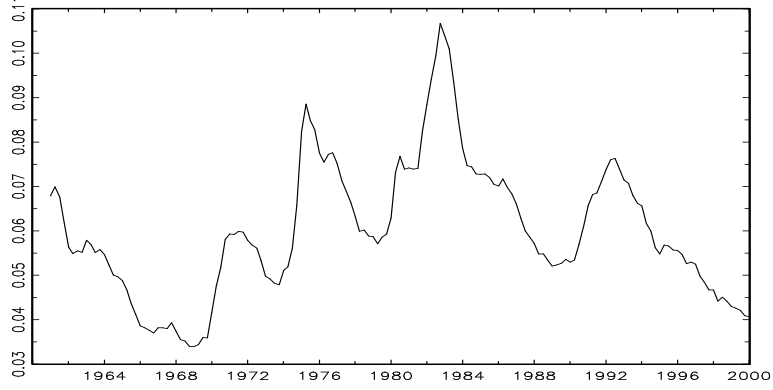


Figure 5: Unemployment Rate of the U.S.

NAIRU more consistent with the actual unemployment rate. The result is

$$\begin{aligned} \pi_t = & 0.0035 + 1.594\pi_{t-1} - 0.832\pi_{t-2} \\ & \quad (0.001) \quad (6.623) \quad (4.961) \\ & - 0.247\text{food}_{t-1} + 0.322\text{food}_{t-2} - 0.017\text{fuel}_{t-1} \\ & \quad (1.722) \quad (2.255) \quad (1.398) \\ & + 0.030\text{fuel}_{t-2} + 0.181r_t - 0.304(U_t - U_t^N) + e_t. \\ & \quad (1.599) \quad (1.532) \quad (0.902) \end{aligned}$$

The estimate of the standard deviation of e_t is 0.010, with t-statistic being 8.982, and the variance of η_t is assumed to be 2.6×10^{-6} . The estimated NAIRU of Italy is presented in Figure 4D.

We also undertake the estimation for the U.S. and estimate the NAIRU with and without “supply shocks” for 1962:3-1999:4.

In the estimation without supply shocks, only four lags of the inflation rate and unemployment gap are included in the regression and the result is

$$\begin{aligned} \pi_t = & 0.002 + 1.321\pi_{t-1} - 0.243\pi_{t-2} \\ & \quad (0.002) \quad (16.045) \quad (2.199) \\ & - 0.121\pi_{t-3} + 0.015\pi_{t-4} - 0.065(U_t - U_t^N) + e_t. \\ & \quad (61.971) \quad (0.350) \quad (8.864) \end{aligned}$$

The estimate of the standard deviation of e_t is 0.004, with t-statistic being 15.651, and the variance of η_t is predetermined as 4.5×10^{-6} . The unemployment rate of the U.S. is presented in Figure ?? and the estimated NAIRU without supply shocks is presented in Figure 4E, very similar to the result of Gordon (1997).

Considering supply shocks which include price changes in food, energy and imports, we have the following result for the U.S.:

$$\begin{aligned} \pi_t = & 0.002 + \underset{(0.091)}{0.957} \pi_{t-1} - \underset{(1.350)}{0.151} \pi_{t-2} \\ & - \underset{(0.647)}{0.070} \pi_{t-3} + \underset{(1.822)}{0.120} \pi_{t-4} + \underset{(4.547)}{0.062} food_t \\ & + \underset{(0.468)}{0.007} fuel_{t-1} + \underset{(3.808)}{0.025} im_{t-1} - \underset{(3.036)}{0.060} (U_t - U_t^N) + e_t. \end{aligned}$$

The estimate of the standard deviation of e_t is 0.003, with t-statistic being 16.217, and the variance of η_t is predetermined as 4×10^{-6} . The estimated NAIRU with supply shocks is presented in Figure 4F.

Up to now, we have shown evidence for time varying reaction coefficients to the Phillips curve and the time varying NAIRU, both taking account of the Lucas critique. The reason of why we estimate the time varying α_{ut} will be more clear in the next section, where we show that the optimal monetary reaction function is related to the coefficients in the Phillips curve. In our model we show that changing coefficients in the Phillips are indeed related to time varying monetary policy reaction.

3 Time Varying Monetary Policy Reaction Function

Our results on the estimation of the time varying NAIRU and parameters in the Phillips curve suggest the further question of how this may be related to time varying monetary policy. We first derive an optimal monetary policy rule from a simple linear quadratic (LQ) control model and then investigate why and how the monetary policy rule may be state dependent.

3.1 Derivation of the Optimal Monetary Policy Rule from A Simple Model

As many other papers on monetary policy, see Svensson (1997 and 1998), Beck and Wieland (2001) and Clarida, Gali and Gertler (1999) for example, we assume that central banks pursue a monetary policy by minimizing a quadratic loss function such as

$$\sum_{t=0}^{\infty} \rho^t U_t, \quad U_t = (\pi_t - \pi^*)^2 + \lambda y_t^2, \quad \lambda > 0,$$

subject to

$$\pi_{t+1} = \alpha_1 \pi_t + \alpha_2 y_t, \tag{5}$$

$$y_{t+1} = \beta_1 y_t - \beta_2 \{(R_t - \pi_t) - \bar{R}\}, \quad \beta_2 > 0, \tag{6}$$

where ρ is the discount rate bounded between 0 and 1, y_t the output gap, π_t the actual inflation rate, π^* the inflation target which will be normalized to zero in our model, R_t the short-term interest rate, \bar{R} the long run equilibrium interest rate and λ is the relative weight put on output gap stabilization by the central bank.

Eq.(5) is the so-called Phillips curve and (6) the IS curve. The reason why the output gap is used here in the Phillips curve, instead of unemployment gap, is that we want to derive a monetary policy rule comparable with the Taylor rule (Taylor 1993), which has become popular in monetary policy studies. The relation between unemployment rate and output can be described by the Okun's Law, which states that when output rises, the unemployment rate falls, and when output falls, the unemployment rate rises. There seems to exist a negative relation between the two coefficients α_{ut} in eq.(1) and α_2 in eq.(5), which can be simplified as

$$\alpha_{ut} = c\alpha_2, \quad c < 0, \quad (7)$$

where c is a constant.

To solve this LQ control problem, we can solve a simpler one first by ignoring the constraint on y_{t+1} at the moment,⁴ as long as there exists a unique solution which minimizes the objective function. This is true in our model, since the return function is quadratic and the constraints are linear. The simpler problem is

$$V(\pi_t) = \underset{y_t}{Min} [(\pi_t^2 + \lambda y_t^2) + \rho V(\pi_{t+1})] \quad (8)$$

subject to

$$\pi_{t+1} = \alpha_1 \pi_t + \alpha_2 y_t. \quad (9)$$

Equation (8) is the so-called Bellman equation and $V(\pi_t)$ is the value function, with y_t being the control variable now. For the linear-quadratic problem like ours, we know that the value function must also be quadratic. Therefore, the value function can be assumed as

$$V(\pi_t) = \Omega_0 + \Omega_1 \pi_t^2, \quad (10)$$

where Ω_0 and Ω_1 remain to be determined (we need only to determine Ω_1 since the goal is to derive the interest rate rule, not to get the optimal value of the objective function). The first-order condition turns out to be

$$\lambda y_t + \rho \alpha_2 \Omega_1 \pi_{t+1} = 0,$$

from which we get

$$\pi_{t+1} = -\frac{\lambda}{\rho \alpha_2 \Omega_1} y_t. \quad (11)$$

Substituting (5) into (11) gives

$$y_t = -\frac{\rho \alpha_1 \alpha_2 \Omega_1}{\lambda + \rho \alpha_2^2 \Omega_1} \pi_t, \quad (12)$$

⁴see also the appendices in Svensson (1997 and 1998).

and after substituting this equation back into (5), we have

$$\pi_{t+1} = \frac{\alpha_1 \lambda}{\lambda + \rho \alpha_2^2 \Omega_1} \pi_t. \quad (13)$$

By applying (8), (10) and (11), the envelop theorem gives us the following equation

$$V_\pi(\pi_t) = 2 \left(1 + \frac{\alpha_1^2 \rho \lambda \Omega_1}{\lambda + \rho \alpha_2^2 \Omega_1} \right) \pi_t,$$

and from (10), we know that

$$V_\pi(\pi_t) = 2\Omega_1 \pi_t,$$

these two equations tell us that

$$\Omega_1 = 1 + \frac{\alpha_1^2 \rho \lambda \Omega_1}{\lambda + \rho \alpha_2^2 \Omega_1}.$$

The right-hand side of this equation has limit $1 + \frac{\alpha_1^2 \lambda}{\alpha_2^2}$ as $\Omega_1 \rightarrow \infty$. The root of Ω_1 larger than one can therefore be solved from the equation

$$\Omega_1^2 - \left[1 - \frac{(1 - \rho \alpha_1^2) \lambda}{\rho \alpha_2^2} \right] \Omega_1 - \frac{\lambda}{\rho \alpha_2^2} = 0,$$

which gives the solution of Ω_1 :

$$\Omega_1 = \frac{1}{2} \left(1 - \frac{\lambda(1 - \rho \alpha_1^2)}{\rho \alpha_2^2} + \sqrt{\left(1 - \frac{\lambda(1 - \rho \alpha_1^2)}{\rho \alpha_2^2} \right)^2 + \frac{4\lambda}{\rho \alpha_2^2}} \right). \quad (14)$$

By substituting $t + 1$ for t into (12), we have

$$y_{t+1} = -\frac{\rho \alpha_1 \alpha_2 \Omega_1}{\lambda + \rho \alpha_2^2 \Omega_1} \pi_{t+1}. \quad (15)$$

Substituting π_{t+1} and y_{t+1} into equation (5) and then (6) into (15) with some computation, we get the optimal decision rule for the short-term interest rate:

$$R_t = \bar{R} + \left(1 + \frac{\rho \alpha_1^2 \alpha_2 \Omega_1}{(\lambda + \rho \alpha_2^2 \Omega_1) \beta_2} \right) \pi_t + \left(\frac{\rho \alpha_2^2 \alpha_1 \Omega_1}{\beta_2 (\rho \alpha_2^2 \Omega_1 + \lambda)} + \frac{\beta_1}{\beta_2} \right) y_t. \quad (16)$$

This is similar to the simple Taylor rule (Taylor 1993). Let

$$\beta_\pi = 1 + \frac{\rho \alpha_1^2 \alpha_2 \Omega_1}{(\lambda + \rho \alpha_2^2 \Omega_1) \beta_2},$$

$$\beta_y = \frac{\rho \alpha_2^2 \alpha_1 \Omega_1}{\beta_2 (\rho \alpha_2^2 \Omega_1 + \lambda)} + \frac{\beta_1}{\beta_2},$$

we have

$$R_t = \bar{R} + \beta_\pi \pi_t + \beta_y y_t. \quad (17)$$

As can be seen from the equations above, β_π and β_y depend on the following parameters: ρ , α_1 , α_2 , λ , β_1 and β_2 . They will change as long as one of these parameters changes. So, it seems unreasonable to assume that β_π and β_y are fixed for all times and it is therefore necessary for the central bank to know how these coefficients may change with the changes of economic conditions. This is what we subsequently focus on. As mentioned above, as long as one of the mentioned parameters (α_2 for instance) changes, β_π and β_y will change. The change of α_{ut} has been explored in Section 2 and, as a result of eq.(7), it is obvious that α_2 can also be time varying, which in turn leads to time varying β_π and β_y in (17).

3.2 Estimates of the Time Varying Taylor Rule

In the previous subsection, we have presented theoretical background of why monetary reaction function may be state-dependent. Here we explore the evidence of how β_π and β_y may change in the monetary policy rule, or in other words, in the simple Taylor rule. We again resort to the Maximum Likelihood estimation by way of the Kalman filter. The corresponding state-space model of our problem is defined as follows:⁵

$$y_t = \beta_t' x_t + \epsilon_t, \quad (18)$$

$$\beta_t - \bar{\beta} = \phi(\beta_{t-1} - \bar{\beta}) + \eta_t, \quad (19)$$

with

$$y_t = R_t, \quad x_t = \begin{pmatrix} 1 \\ \pi_t - \pi^* \\ \hat{y}_t \end{pmatrix}, \quad \beta_t = \begin{pmatrix} \beta_{ct} \\ \beta_{\pi t} \\ \beta_{yt} \end{pmatrix},$$

with \hat{y}_t being the output gap. The disturbances ϵ_t and η_t have Gaussian distributions with zero mean and constant variance.

We present the empirical results for Germany and then jointly for France, Italy and the U.K. below. The data source is the same as in section 2. The output gap is measured by the percent deviation of Industrial Production Index from its HP filtered-trend. The inflation targets for Germany, France, the U.K. and Italy are assumed to be 0.02, 0.025, 0.025 and 0.03 respectively.⁶ The short-term interest rates of the above countries are measured by 3-month FIBOR, 3-month PIBOR, 3-month interbank loans and official discount rate respectively.

The German data from 1960-1998 generate the ϕ as $\begin{pmatrix} 0.935 & 0 & 0 \\ 0 & 0.892 & 0 \\ 0 & 0 & 0.925 \end{pmatrix}$.

All the elements of ϕ are smaller than 1 in absolute value, so the coefficients are

⁵Here we use the so-called Return-to-normality model to estimate the paths of the time-varying coefficients. The details of the state space model and the Kalman filter is presented in the appendix of Semmler, Greiner and Zhang (2001).

⁶We refer to Clarida, Gali and Gertler (1998) for the inflation targets of the countries studied. A brief justification can be found in Semmler, Greiner and Zhang (2001).

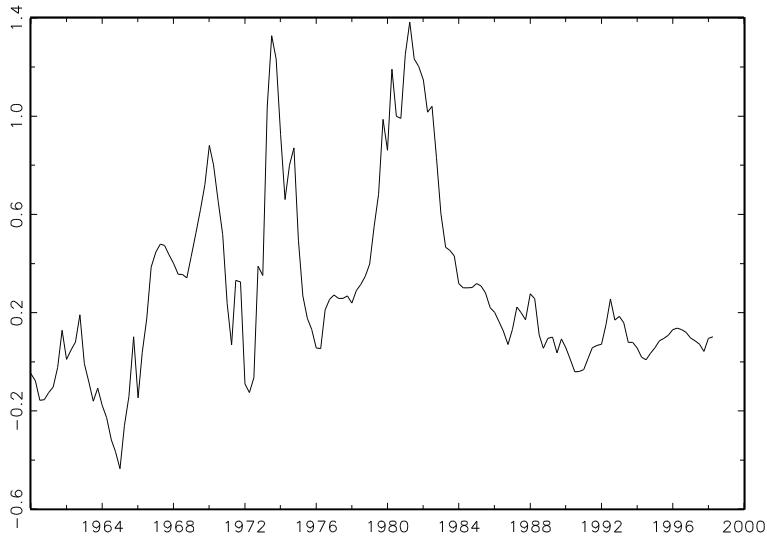


Figure 6: $\beta_{\pi t}$ for Germany

stationary, and we get the $\bar{\beta}$ as $\begin{pmatrix} 0.052 \\ 0.260 \\ 0.294 \end{pmatrix}$, which means that β_{ct} evolves around 0.052, $\beta_{\pi t}$ around 0.260 and β_{yt} around 0.294. The paths of $\beta_{\pi t}$ and β_{yt} are presented in Figure 6-7.⁷

As shown in Figure 6, $\beta_{\pi t}$ experiences very significant changes. On the whole, the switching of $\beta_{\pi t}$ is similar to that of $\pi_t - \pi^*$, except in the 1960s.⁸ This means that when the inflation rate is high, $\beta_{\pi t}$ is also high, and vice versa (see Figure 2 for Germany inflation rate). In 1970, 1974 and 1981, $\beta_{\pi t}$ reaches some peaks, when the interest rate and inflation rate were also at their peaks. In the 1960s, we find that $\beta_{\pi t}$ and $\pi_t - \pi^*$ evolves in opposite directions most of the time, especially during 1965-1970. From 1960-1965, $\beta_{\pi t}$ is below zero most of the time. There is a significant structural change around 1979 and a small change around 1989 for $\beta_{\pi t}$.

Since the switching of $\beta_{\pi t}$ is similar to that of $\beta_t - \bar{\beta}$ most of the time, we can presume that the changes of the coefficients are caused by the corresponding variable to some extent, that's to say, they are state-dependent. A more direct evidence is to regress the coefficients on the correspondent variable, that is to run an OLS regression for the following equation

$$\beta_{\pi t} = c_1 + c_2(\pi_t - \pi^*) \quad (20)$$

⁷We will not present the paths of β_c below, since we are not interested in its changes.

⁸The result that β_{π} and inflation gap evolve differently at the beginning of the sample period for Germany may be related to initial startup idiosyncracies that may be associated with kicking off the Kalman filter algorithm.

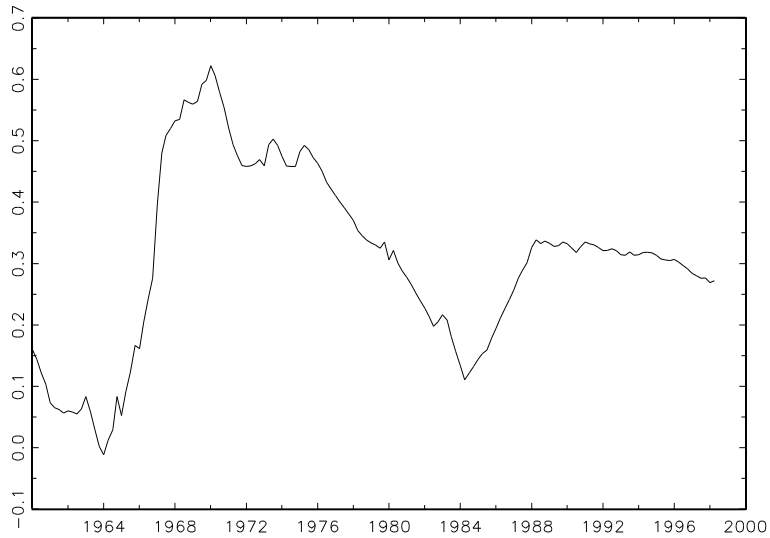


Figure 7: β_{yt} for Germany

Period	c_1		c_2		R^2
	Estimate	T-st.	Estimate	T-st.	
1960:1-1964:4	-0.124	3.663	9.567	2.160	0.206
1965:1-1969:4	0.360	6.684	-21.880	4.408	0.519
1970:1-1979:4	0.165	1.379	9.342	2.571	0.148
1980:1-1989:4	0.350	9.700	17.320	10.912	0.758
1990:1-1993:4	0.033	1.295	3.038	2.593	0.324
1994:1-1998:2	0.069	8.579	-4.947	3.368	0.415
1960:1-1998:2	0.157	1.200	10.381	7.435	0.267

Table 1: Results of eq.(20) (Germany)

For the whole period 1960-1998, we have

$$\beta_{\pi t} = 0.157 + 10.381(\pi_t - \pi^*)$$

with $R^2=0.267$ and t-statistics for c_2 being 7.435. The R^2 is not large enough to confirm our statement above. But we will get better results if we divide the whole period into subperiods for about every five or ten years, or shorter time if there are break-points. The results are presented in Table 1, where we have higher R^2 and significant t-statistics. Note that the R^2 ranges between 0.148 and 0.758, which indicates that for some subperiods, there may exist some other variables responsible for the change of $\beta_{\pi t}$.

As shown in Figure 7, β_{yt} also experiences changes, but it is less drastic than $\beta_{\pi t}$. It attained its highest level around 1970 and then continued to decrease until the middle of the 1980s and remains at a relatively high level during the

Period	c_1		c_2		R^2
	Estimate	T-st.	Estimate	T-st.	
1960:1-1969:4	0.217	6.787	-9.670	2.549	0.146
1970:1-1974:4	0.476	33.101	5.019	2.275	0.223
1975:1-1979:4	0.387	53.664	-6.012	7.082	0.736
1980:1-1984:4	0.224	18.191	5.800	2.947	0.326
1985:1-1989:4	0.260	17.848	-10.884	2.256	0.220
1990:1-1994:4	0.322	249.520	0.599	3.774	0.504
1995:1-1998:2	0.297	80.831	-2.171	2.132	0.221
1960:1-1998:2	0.306	25.271	-1.563	0.892	0.005

Table 2: Results of eq.(21) (Germany)

1990s. We also undertook the OLS regression for the equation

$$\beta_{yt} = c_1 + c_2 \hat{y}_t \quad (21)$$

and present the result in Table 2, where we find that t-statistics are usually significant, but the R^2 changes between 0.146 and 0.736 for the subperiods and it is only 0.005 for the whole period. Although β_{yt} appears to be state dependent, there may also be other variables responsible for the change of β_{yt} besides the output gap.

Next, we compare the monetary policies for the so-called E3 countries (France, the U.K. and Italy),⁹ whose central banks have less control over the domestic monetary policy than those of the so-called G3 countries (Germany, Japan and the U.S.). We may find something interesting for these three countries. As we will see below, the changes of the response coefficients for these three countries have some similarities. We will analyze this problem briefly. $\beta_{\pi t}$ of the three countries are presented in Figure 8 (we present $\beta_{\pi t}$ of the U.K. from 1970 on, in order to be consistent with those of the other two countries).

Figure 8 shows that the $\beta_{\pi t}$ of the three countries experience some significant changes in the 1970s and then remained at a relatively stable and high level from the middle of the 1980s on. That means the inflation deviation plays an important role in the three countries' policy making after 1980. Moreover, the changes of $\beta_{\pi t}$ for the U.K. and France look similar before 1985, though the $\beta_{\pi t}$ of France stayed at a higher level than that of the U.K. in this period.

The inflation rates of the three countries can be seen in Figure 2, which shows that the inflation rates of the three countries also experience some similar changes, as mentioned in subsection 2.1. The similarity of inflation rates between the three countries leads to some similarity of $\beta_{\pi t}$. The β_{yt} of the three countries are presented in Figure 9.

The changes of β_{yt} for Italy and France are also similar most of the time, both of them decreased to the lowest level between 1992 and 1993, when the

⁹see Clarida, Gali and Gertler (1998).

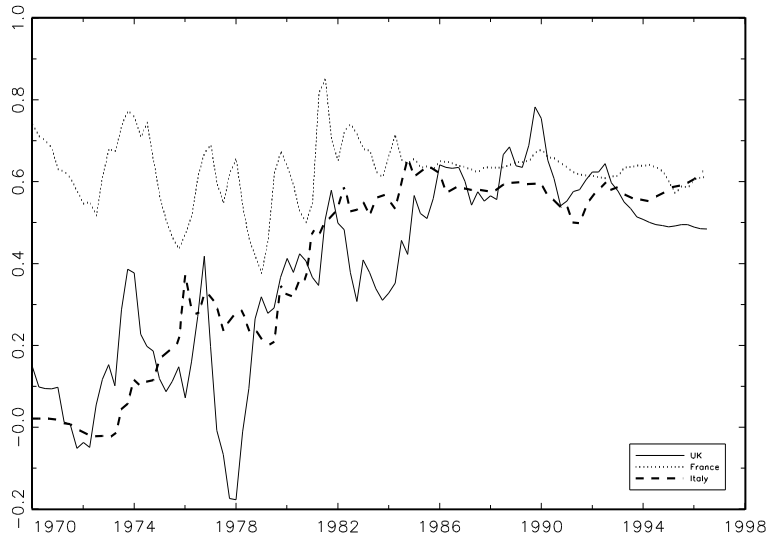


Figure 8: $\beta_{\pi t}$ of France, Italy and the U.K.

crisis of the EMS occurred. The β_{yt} of U.K. is very smooth. In Figure 10, we present the output gaps of the three countries.

Figure 10 shows that the output gaps of the three countries also experience similar changes. From the evidence above, we would like to say that there is some consistency between the monetary policies of the E3 countries. We can observe that for all three countries the response coefficient of the inflation gap moved up and stayed high in the 1990s and that the response coefficient for the output gap is almost constant except when Germany raised the interest rate after the German unification. The other countries had to raise the interest rate too, in spite of a negative output gap (see β_{yt} for France and Italy).

We also estimate $\beta_{\pi t}$ and β_{yt} for the U.S., which are presented in Figure 11-12. The short-term interest rate of the U.S. is measured by the Federal funds rate. The estimates of ϕ and $\bar{\beta}$ are $\begin{pmatrix} 0.991 & 0 & 0 \\ 0 & 0.893 & 0 \\ 0 & 0 & 0.674 \end{pmatrix}$ and $\begin{pmatrix} 0.050 \\ 0.448 \\ 0.705 \end{pmatrix}$ respectively. The paths of $\beta_{\pi t}$ and β_{yt} will play an important role in the simulation in the next section.

Overall, the results of the time varying coefficient estimations favor the statement that the coefficients of the economic variables in monetary policy rules are not fixed but time varying. In fact they change with the economic environment and are thus state dependent. Woodford (2001) undertakes a welfare loss analysis. There it is shown that there is certain relation between the coefficients in the monetary reaction function and the relative weight λ in the objective function. Therefore, it would also be interesting if we relax the assumption of a fixed λ in the central bank's objective function. This is discussed in Greiner and Semmler (2001).

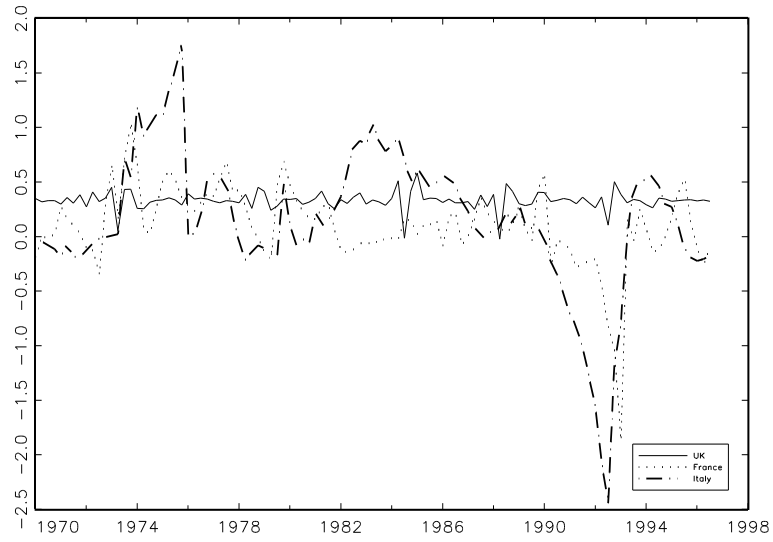


Figure 9: β_{yt} of France, Italy and the U.K.

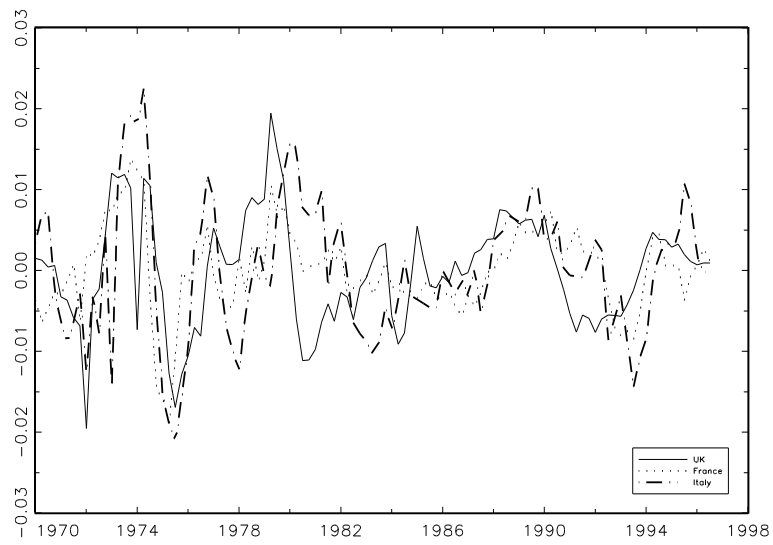


Figure 10: Output Gaps of France, Italy and the U.K.

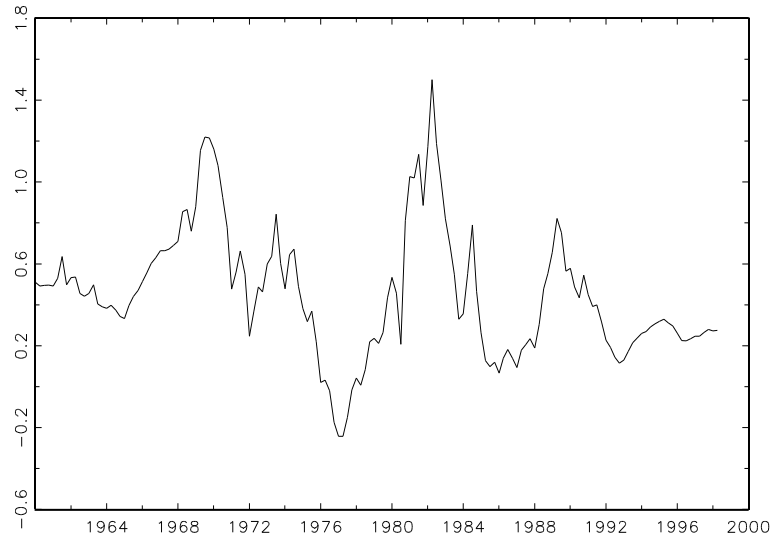


Figure 11: $\beta_{\pi t}$ for the U.S.

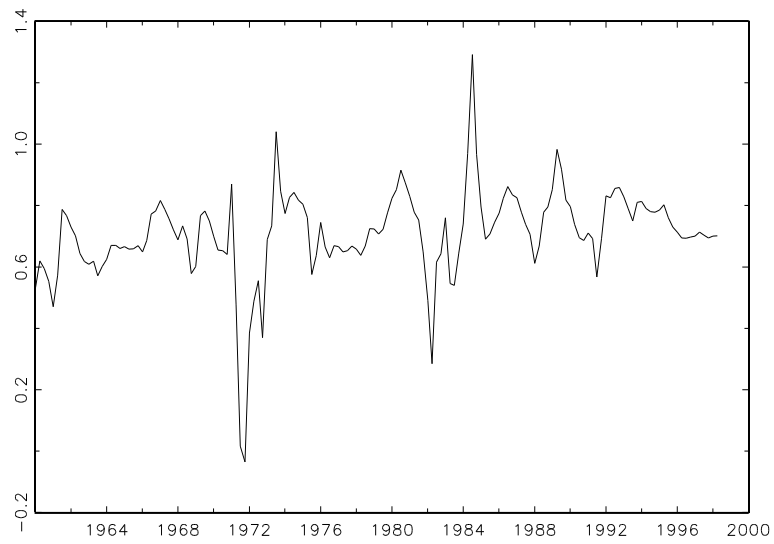


Figure 12: β_{yt} for the U.S.

4 Euro-Area Monetary Policy Effects Using the U.S. Monetary Rules

It is well known that the economy of EU has been in a worse situation than that of the U.S. in the 1990s, which is reflected in their GDP growth rates and unemployment rates respectively. The differences in the GDP growth and unemployment performance of the EU and the U.S. may be seen to have also been influenced by differences in monetary policies. Similar to Peersman and Smets (1998), we use the German call money rate to measure the monetary policy in the EURO-area.¹⁰ The aggregate inflation rate and output gap of the EURO-area are measured respectively as the GDP weighted sums of the inflation rates and output gaps of the three countries Germany, France and Italy. Similar aggregation of the data can also be found in Taylor (1999a). The short-term interest rate of the U.S. is measured by the Federal funds rate. The data source is the same as those of the past two sections.¹¹ Difference in the interest rates can be seen in Figure 13. Before 1994, the interest rate of the U.S. is much lower (4.9 percent on average) than that of the EURO-area (8.47 percent on average), and after 1994, the EURO-area implemented a looser monetary policy (4.01 percent on average) than the U.S. (5.25 percent on average). For the whole period of 1990-98, the average rate of the U.S. is 5.14 percent, while that of EURO-area is 6.07 percent. Thus an interesting question is what would have happened if the EURO-area had followed the monetary policy rule of the U.S. in the 1990s. Next we attempt to answer this question. On the basis of the theoretical and empirical research in the previous two sections, we will simulate the inflation rate and output gap for the EURO-area for 1990-98, assuming that the EURO-area had followed the monetary policy of the U.S.. The first subsection undertakes simulations under the assumption that the Taylor rule has time varying response coefficients and the second subsection assumes that the Taylor rule has fixed coefficients. A similar counterfactual study has been undertaken by Taylor (1999b) using the pre-Volcker policy interest rate reaction function to study the macroeconomic performance of the Volcker and post-Volcker periods. Therefore, in the third subsection, we undertake the simulations with the two so-called suggested Taylor rules.

4.1 Simulation Using the Time Varying Coefficient Taylor Rule of the U.S.

The Taylor rule with time varying response coefficients can be presented as:

$$R_t = \bar{R} + \beta_{\pi t}(\pi_t - \pi^*) + \beta_{yt}y_t, \quad (22)$$

where \bar{R} is the long-run equilibrium interest rate, and other variables are the same as in section 3 (y_t is the output gap now, while the output gap is indicated

¹⁰Peersman and Smets (1998) justified briefly why the Germany day-to-day rate can be taken as the short-term interest rate of the EURO-economy.

¹¹The GDP weights are computed as follows: We first computed the GDP value of the three countries in US dollar, and then compute the weights of the three countries.

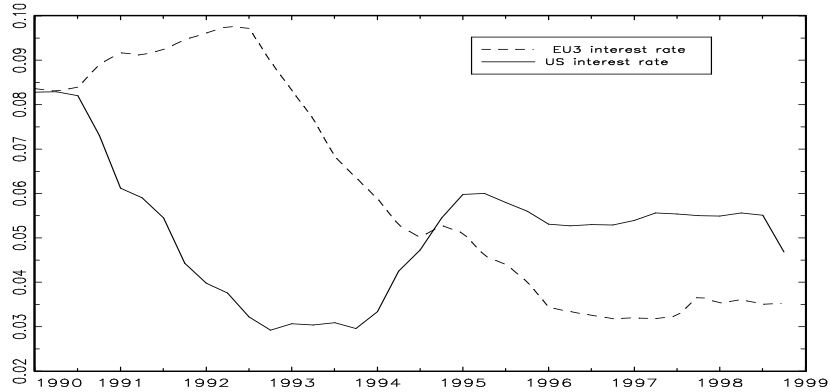


Figure 13: Interest Rates of the U.S. and the EURO-Area

by \hat{y}_t in section 3). The paths of $\beta_{\pi t}$ and $\beta_{y t}$ of the U.S. are presented in Figure 11 and 12.

Next, we assume that the EURO-area follows the monetary policy of the U.S. by determining the interest rate using the HP filter trends of $\beta_{\pi t}$ and $\beta_{y t}$ of the U.S. instead of the exact paths of $\beta_{\pi t}$ and $\beta_{y t}$, since it seems unreasonable to assume that the EURO-area would have followed exactly the same rule as the U.S.. We simulate the EU3 interest rate with equation (22) by defining \bar{R} as the average real interest rate of the EURO-area in the 1990s and substituting the U.S. $\beta_{\pi t}$ and $\beta_{y t}$ trends for $\beta_{\pi t}$ and $\beta_{y t}$. The inflation target is assumed to be as 2.5 percent. The simulated EURO-area interest rate (denoted by simulated interest rate in Figure 14) is presented in Figure 14 together with the actual EURO-area interest rate. The simulated rate is much lower (3.56 percent on average before 1995) than the actual rate in the first half of the 1990s and close to the actual rate after 1994. The average simulated rate is 3.39 percent for 1990-98. The simulations of the EURO-area inflation rate and output gap will be undertaken from the IS-Phillips curves:¹²

$$\pi_t = a_1 + a_2\pi_{t-1} + a_3\pi_{t-2} + a_4\pi_{t-3} + a_5y_{t-1}, \quad (23)$$

$$y_t = b_1 + b_2y_{t-1} + b_3y_{t-2} + b_4(R_{t-1} - \pi_{t-1}), \quad (24)$$

where all the variables have the same interpretations as in equation (22). In order to simulate the π_t and y_t from equation (23)-(24), we need to know the values of the coefficients, a_i and b_i ($i = 1...4$), which will be generated by the

¹²The numbers of lags included depend on the t-statistics of the OLS estimated coefficients of the corresponding lags, if the t-statistics of the coefficients are insignificant, we exclude the corresponding lags. This backward-looking model is justified in Rudebusch and Svensson (1999).

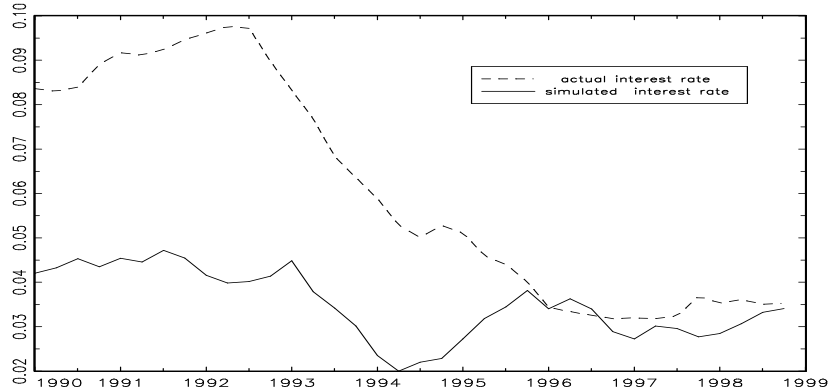


Figure 14: Actual and Simulated Interest Rate of the EURO-Area

(SUR) estimation of equation (23)-(24) with actual data for 1990-98. The results for this system are as follows, with t-statistics in parentheses,¹³

$$\begin{aligned} \pi_t &= -0.0006 + 0.984\pi_{t-1} - 0.291\pi_{t-2} + 0.286\pi_{t-3} + 0.149y_{t-1} \\ &\quad (0.254) \quad (6.165) \quad (1.307) \quad (1.743) \quad (1.014) \\ R^2 &= 0.852 \\ y_t &= 0.0002 + 1.229y_{t-1} - 0.403y_{t-2} - 0.008(R_{t-1} - \pi_{t-1}) \\ &\quad (0.277) \quad (7.841) \quad (2.467) \quad (0.334) \\ R^2 &= 0.799. \end{aligned}$$

The residual covariance is 6.82×10^{-11} . After substituting the simulated interest rate into these equations, we have the simulations of π_t and y_t .¹⁴ The simulated output gap is presented in Figure 15. We find that the simulated

¹³The t-statistics of the last terms of these two equations are unfortunately insignificant. The results seem sensitive to the period we cover and how potential output is computed. If we use a linear quadratic trend of log industrial production index as potential output, for example, we have the following results for the data 1986-98:

$$\begin{aligned} \pi_t &= 0.004 + 1.117\pi_{t-1} - 0.341\pi_{t-2} + 0.072\pi_{t-3} + 0.184y_{t-1} \\ &\quad (1.806) \quad (8.150) \quad (1.719) \quad (0.556) \quad (1.934) \\ R^2 &= 0.830 \\ y_t &= 0.001 + 1.254y_{t-1} - 0.275y_{t-2} - 0.045(R_{t-1} - \pi_{t-1}) \\ &\quad (1.772) \quad (9.376) \quad (1.835) \quad (1.714) \\ R^2 &= 0.909 \end{aligned}$$

with determinant residual covariance being 6.44×10^{-11} . The t-statistics of the last terms are now more significant. Since the t-statistics significance of these terms has little effect on our simulations, we do not discuss how output gap should be defined here.

¹⁴As for the 3 initial lags of inflation and 2 initial lags of output gap, we just take the actual inflation rate data from 1989.2-89.4 and actual output gap data from 1989.3-89.4.

output gap declines at the beginning of the 1990s and increases a little in 1994. The simulated and actual output gaps are presented in Figure 16. Unlike the actual output gap, which experiences significant decreases during 1992-94 and 1995-97, the simulated output gap is always positive and smoother than the actual one. The simulated and actual inflation rates are presented in Figure 17. We find that the simulated inflation is almost linear and is lower than the actual inflation rate most of the time.

4.2 Simulation Using the Fixed Coefficient Taylor Rule of the U.S.

In the last subsection, we have simulated the inflation rate and output gap of the EURO-area under the assumption that the coefficients of output gap and inflation in the Taylor rule are time varying. In this subsection, we assume that the coefficients are fixed. Using the U.S. data of 1990.1-98.4, we get the estimation (t-statistics in parentheses)

$$R_t = 0.049 + 0.735(\pi_t - \pi^*) + 0.703y_t, \quad R^2 = 0.520.$$

(26.970)
(5.077)
(5.469)

With these coefficients, we can get the simulated interest rate of the EURO-area, which is presented in Figure 18, together with the actual the EURO-area interest rate and the simulated interest rate from last subsection (simulated interest rate 1), the simulated interest rate from this subsection is denoted by simulated interest rate 2. We find that the simulated interest rate under the assumption of fixed coefficients is higher than the simulated rate with the time varying coefficient assumption most of the time and experiences more significant changes, but it is still much lower than the actual interest rate.

Next, we can get the simulated inflation rate and output gap with the simulated interest rate by way of equation (23)-(24). The simulated inflation rate (simulated inflation 2) is presented here together with the actual inflation rate and the simulated inflation from the last subsection (simulated inflation 1) in Figure 19. Figure 19 shows that the simulated inflation from this subsection is slightly lower than the simulated inflation from the previous subsection, which is consistent with the fact that the simulated interest rate of this subsection is higher than that from the last subsection most of the time. The simulated output gap (simulated output gap 2) is presented in Figure 20, together with the simulated output gap from the last subsection (simulated output gap 1) and the actual output gap. We find that the output gap simulated here is also very smooth and close to the simulation from the last subsection.

4.3 Simulation Using the Suggested Taylor Rules and Actual Interest Rate

In the last two subsections, we have undertaken simulations for the EURO-area with both time varying and fixed β_π and β_y chosen as the estimates from monetary reaction function with the U.S. data for the 1990s. In Taylor (1999b),

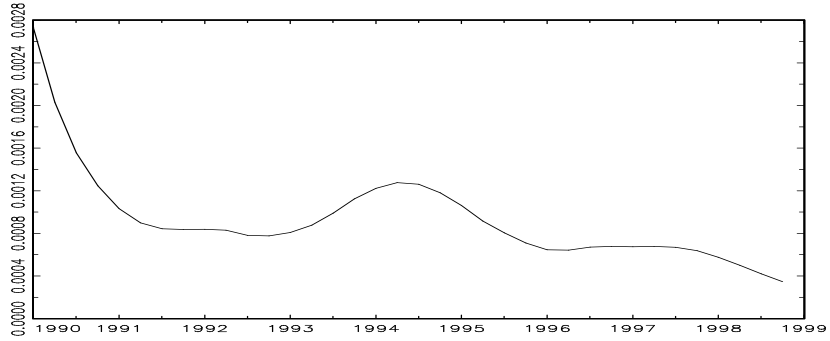


Figure 15: Simulated Output Gap of the EURO-Area

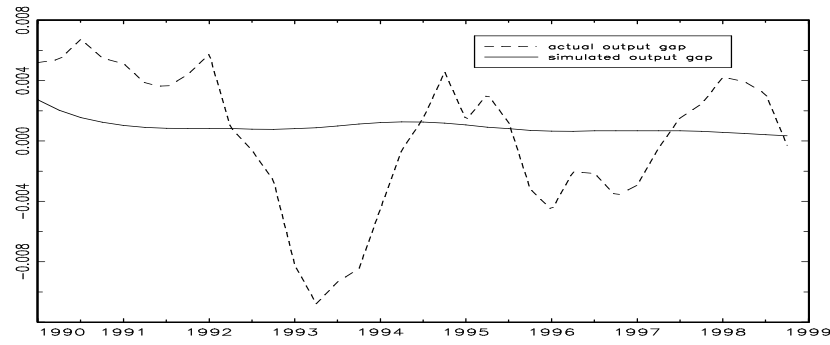


Figure 16: Actual and Simulated Output Gaps of the EURO-Area

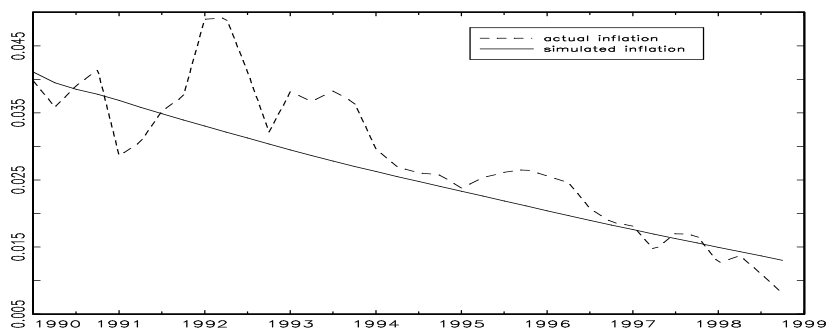


Figure 17: Actual and Simulated Inflation Rates of the EURO-Area

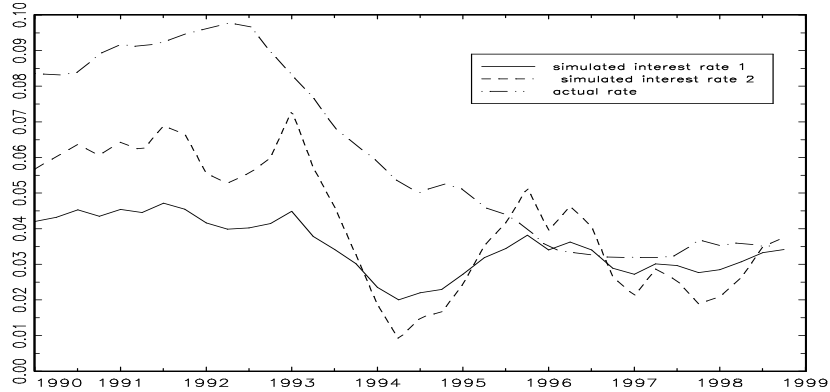


Figure 18: Actual and Two Simulated Interest Rates of the EURO-Area

however, a counterfactual study was undertaken, using the pre-Volcker policy interest rate reaction function to explore the macroeconomic performance of the post-Volcker period. Two suggested rules can be found there. In this subsection, we will undertake simulations for the EURO-area with the two suggested Taylor rules, and make a comparison with the simulations above.

The first rule, which was first stated in Taylor (1993), suggests β_π to be 1.5 and β_y 0.5. Namely,

$$R_t = \bar{R} + 1.5(\pi_t - \pi^*) + 0.5y_t. \quad (25)$$

The second suggestion keeps β_π at 1.5 but raises β_y to 1.0, namely

$$R_t = \bar{R} + 1.5(\pi_t - \pi^*) + y_t. \quad (26)$$

The interest rates simulated with these two suggested rules are presented in Figure 21, together with the actual interest rate and the interest rates simulated in the previous two subsections (simulated interest rate 1 and simulated interest rate 2). The two rules simulated here (Taylor rule 1 and Taylor rule 2) are close to each other, lower than the interest rates simulated in the last two subsections most of the time and much lower than the actual rate.

The simulated output gaps with the suggested Taylor rules are presented in Figure 22. They are positive all the time. They decrease to the lowest points in 1992 and then increase slowly. In Figure 23, we present the four simulated output gaps together, from which we can see the difference easily.¹⁵

¹⁵In Figure 23, simulated output gap 3 and 4 are the output gaps simulated with the two suggested Taylor rules, the same as in Figure 22. Simulated output gap 1 and 2 are the output gaps simulated with time varying and fixed Taylor rules discussed in 4.1 and 4.2 respectively.

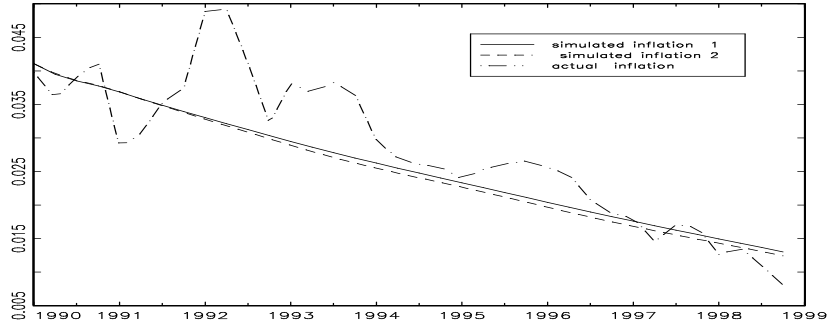


Figure 19: Actual and Simulated Inflation Rates

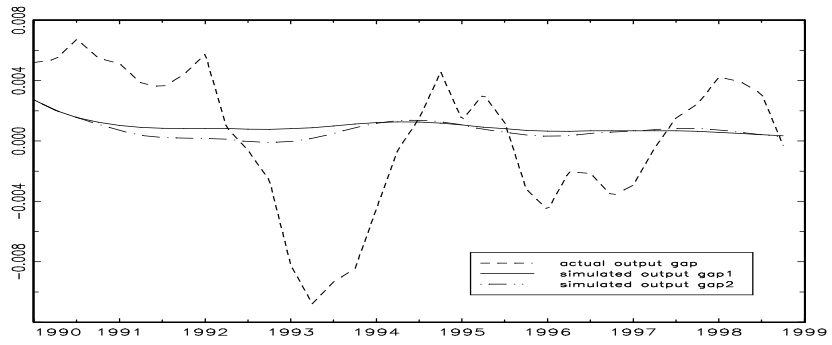


Figure 20: Actual and Simulated Output Gaps

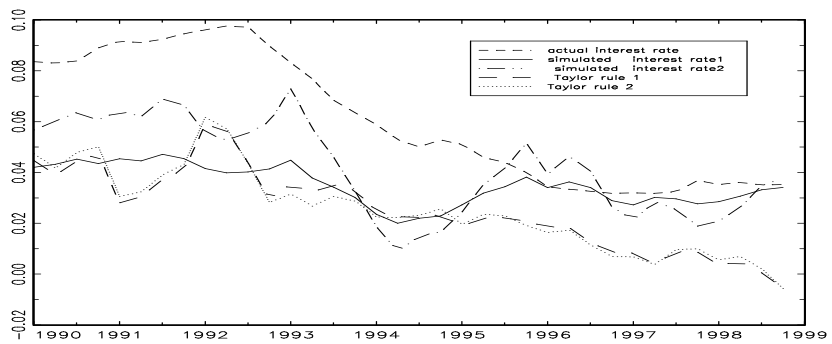


Figure 21: Actual and Simulated Interest Rates

The simulated output gaps with the two suggested Taylor rules are very close to the one simulated with time varying response coefficient interest rate rule before 1995, but after 1995, there is an obvious difference: The former increase slowly, while the latter decreases. The inflation rates simulated with the two suggested Taylor rules (denoted by simulated inflation 3 and 4 respectively in Figure 24) are very close to each other. In Figure 24, we also present the inflation simulations (simulated inflation 1 and 2) from last two subsections. It is very obvious that the four simulations are all similar to a linear trend of the actual inflation rate.

Up to now, we have simulated the inflation rate and output gap for EU3, assuming that different interest rate rules of the U.S. had been followed in the 1990s. The simulations above favor the conclusion that if the EURO-area had followed the interest rate rule of the U.S., the output gap would have been much smaller and smoother than the actual one, while the inflation rate would have been similar to the linear trend of the actual one. One problem to note is, we have used the Germany call money rate as the interest rate for EURO-area, and assume that the actual output and inflation rate are really generated by this interest rate, which was higher than the U.S. rate most of the time in the 1990s. So, the question arising here is, what will happen if simulations are undertaken with the actual rate? Will the output gap simulation really experience drastic changes like the actual one? To answer this question, we will do the simulations with the actual EURO-area rate below. The output gap simulated with the actual EURO-area rate is presented in Figure 25, together with all the other simulated output gaps. It is obvious that the output gap simulated with actual interest rate (denoted by simulated output gap 5 in Figure 25) is lower and experiences more changes than the others. Moreover, like the actual output gap (which experiences a large decrease during 1991 and 1995), it is negative from 1991 to 1996. The simulated inflation with actual EU3 interest rate is presented in Figure 26, denoted by simulated inflation 5, together with other four simulations, it is slightly lower than the other ones.

5 The EURO-Area Monetary Policy and Asset Price Targeting

An interesting feature of the monetary and financial environment in the U.S. as well as the EURO-area in the 1990s was that the inflation rate has decreased and remained relatively stable at a low level, while asset prices—the prices of equities, bonds and foreign exchanges—experienced strong long-run appreciation and depreciation as well as short-term volatility. Some authors have maintained that asset prices should be targeted by monetary authorities only if they affect the inflation and output, see Bernanke and Gertler (2000) for example. But according to other researchers, asset prices should be included as a direct target in central banks' objective function. Such arguments can be found in Dupor (2001), Cecchetti et al (2001) and Semmler and Zhang (2002). In this section,

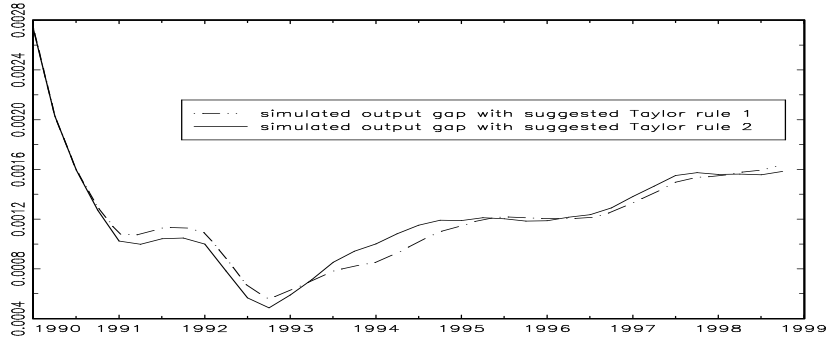


Figure 22: Simulated Output Gaps with Suggested Taylor Rules

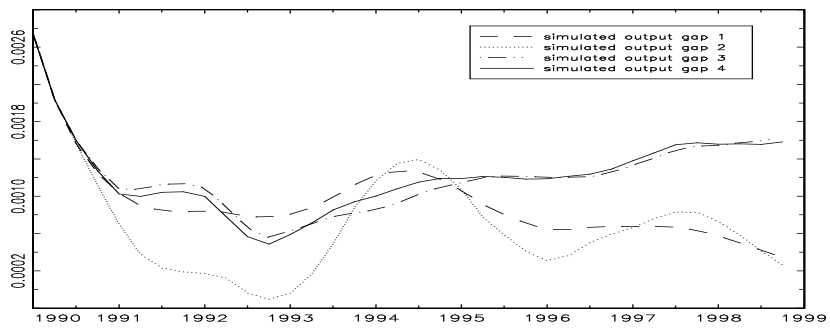


Figure 23: Four Simulated Output Gaps of the EURO-Area

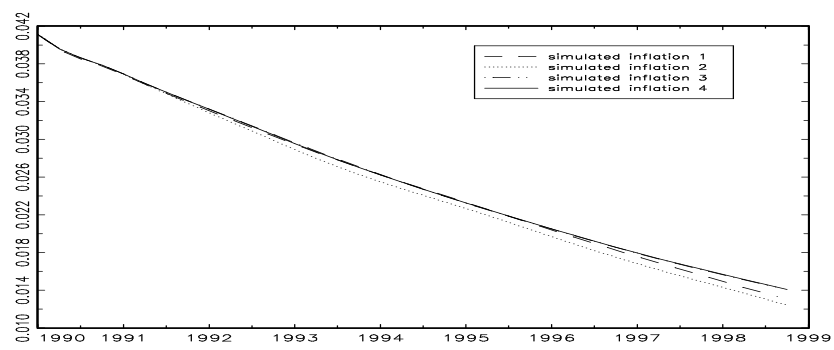


Figure 24: Four Simulated Inflation Rates of the EURO-Area

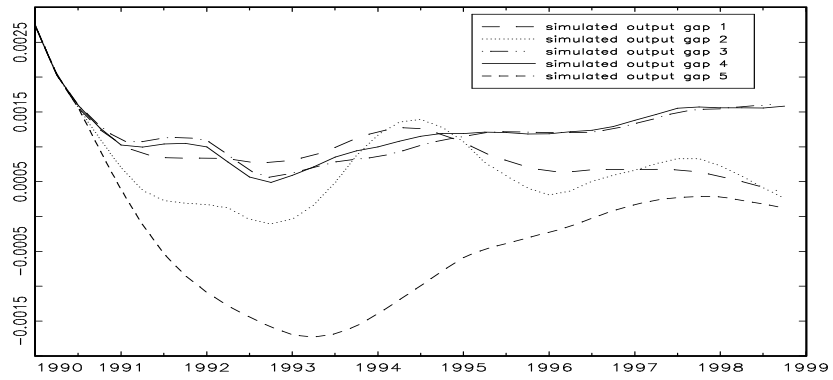


Figure 25: All Simulated Output Gaps of the EURO-Area

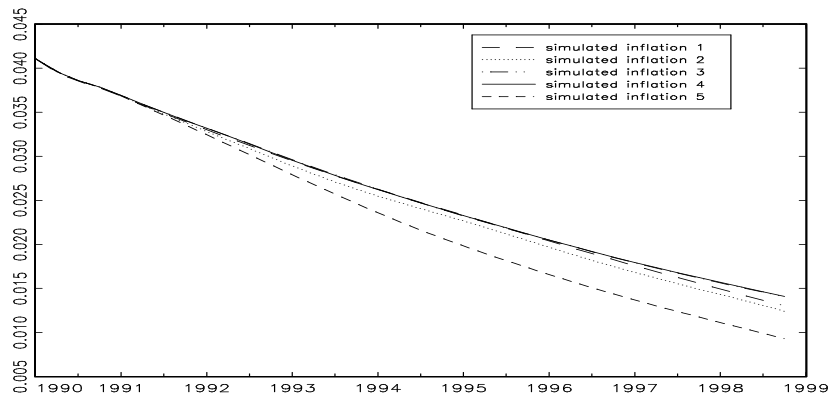


Figure 26: All Simulated Inflation Rates of the EURO-Area

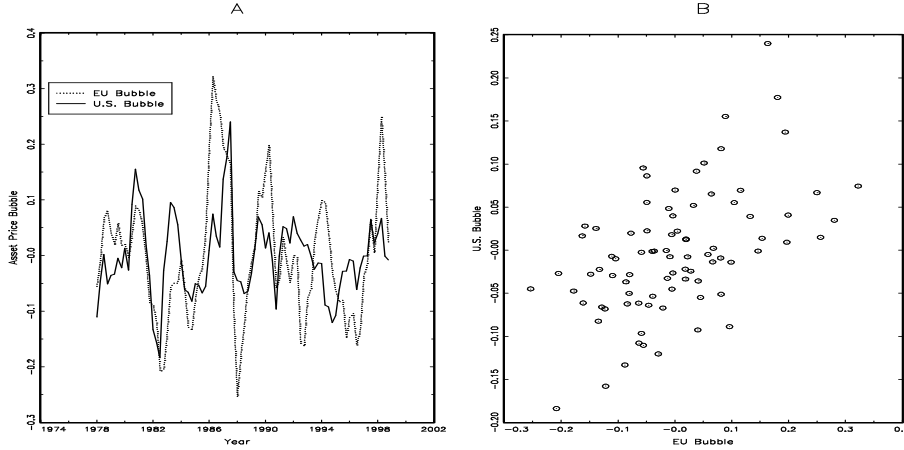


Figure 27: Asset Price Bubble of the EURO-Area and the U.S.

we will explore briefly the relationship of the monetary policy and asset prices for the EURO-area.

5.1 Inflation, Output and Asset Price Bubble

We measure asset prices of the EURO economy with the GDP weighted sum of the Stock Price Index of Germany, France and Italy. The asset price bubble is measured by the percentage deviation of the Stock Price Index from its HP filtered trend. In order to compare the asset price bubble of the EURO-area with that of the U.S., we present the bubbles of the two economies in Figure 27. Figure 27A indicates that the turning points of the asset price bubble of the two economies are similar most of the time. But the asset price bubble of the EURO-area experiences more significant changes with a standard deviation (S.D.) of 0.117 from 1978.1-98.4, while the S.D. of the U.S. asset price bubble for the same period is only 0.072. In Figure 27B we observe an obvious positive correlation between the bubbles of the two economies with a correlation coefficient of 0.534. Moreover, another interesting regression is the following one:

$$B_{eu,t} = 0.002 + 0.724 B_{eu,t-1} + 0.404 B_{us,t}, \quad R^2 = 0.718,$$

(0.233) (11.307) (3.822)

where $B_{eu,t}$ and $B_{us,t}$ stand for asset price bubble of the EURO-area and the U.S. respectively. The regression above also confirms the strong correlation between the asset bubbles of the two economies.

In Figure 28 we present the asset price bubble, inflation rate and output gap of the EURO-area together (note that in Figure 28A we present the fourth lag

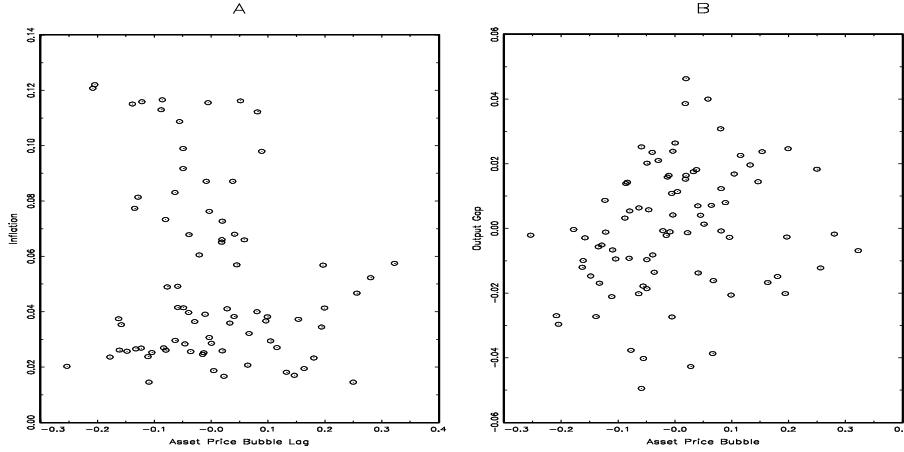


Figure 28: The Asset Price Bubble, Inflation Rate and Output gap of the EURO-Area

of the asset price bubble instead of the current asset price bubble). In Figure 28 we can hardly find a clear relation between the asset price bubble (B_t), inflation and output gap. Next we undertake an OLS regression for these variables. With the EURO-area quarterly data from 1978.1-98.4 we obtain the following results:

$$y_t = 0.0005 + 0.808 y_{t-1} + 0.018 B_t, \quad R^2 = 0.725.$$

(0.410) (14.021) (1.822)

The coefficient of B_t has the right sign with significant t-statistics. For the regression of the inflation rate on the asset price bubble, we have

$$\pi_t = -0.0007 + 0.999 \pi_{t-1} + 0.004 B_t, \quad R^2 = 0.976.$$

(0.620) (55.451) (0.762)

The t-statistic of the coefficient of the asset price bubble is insignificant. This seems to indicate that there exists little correlation between B_t and π_t , but if we try with a fourth lag of B_t in the regression we get a better result (the fourth lag of the asset price bubble and inflation are presented in Figure 28A).

$$\pi_t = -0.0006 + 0.998 \pi_{t-1} + 0.008 B_{t-4}, \quad R^2 = 0.976.$$

(0.518) (55.891) (1.628)

The estimates above seem to indicate that the asset price bubble has to some extent affected (though not greatly) the inflation rate and output gap in the EURO-area economy.¹⁶

¹⁶Further evidence on asset prices, output and inflation is discussed for Japan in Okina et. al. (2000).

5.2 The Asset Price Bubbles and Interest Rate Determination of the EURO-Area

In the last subsection we have explored whether asset price bubble has affected the inflation rate and output gap in the EURO-area. Next, we will explore whether the asset price bubble has been taken into account in the interest determination in the EURO-area. We follow the model of Clarida, Gali and Gertler (1998) (referred to as CGG98 afterwards). This model has also been followed by Smets (1997) and Bernanke and Gertler (2000) to explore the monetary policies in Canada, Australia, the U.S. and Japan.

Smets (1997) follows CGG98 and estimates the monetary reaction function of Canada and Australia by adding three financial variables into the CGG98 model, namely, the nominal trade-weighted exchange rate, ten-year nominal bond yield and a broad stock market index. His conclusion is that the Bank of Canada reduces interest rates significantly in response to an appreciation of the exchange rate, and more surprisingly, changes in the stock market index are also significant in the policy reaction function of the Bank of Canada. For Australia, the coefficients are insignificant.

Bernanke and Gertler (2000) also follow CGG98 by adding stock returns into the model to test whether interest rates respond to stock returns in the U.S. and Japan. Their conclusion is that the Federal Funds rate does not show a significant response to stock returns from 1979-97. For Japan, however, they find different results: For the whole period 1979-97, there is little evidence that the stock market plays a role in the interest rate setting, but for the two subperiods, 1979-89 and 1989-97, the coefficients of stock returns are significant enough (but with different signs).

Next we also follow CGG98 to test whether the EURO-area monetary policy shows a significant response to the stock market. In CGG98, the short-term interest rate is assumed to take the following path:

$$R_t = (1 - \rho)R_t^* + \rho R_{t-1} + v_t, \quad (27)$$

where R_t denotes the short-term interest rate, R_t^* the interest rate target, v_t an i.i.d., and ρ captures the degree of interest rate smoothing. The target interest rate is assumed to be formed in the following way:

$$R_t^* = \bar{R} + \beta(E[\pi_{t+n}|\Omega_t] - \pi^*) + \gamma(E[y_t|\Omega_t] - y_t^*),$$

where \bar{R} is the long-run equilibrium nominal rate, π_{t+n} is the rate of inflation between periods of t and $t+n$, y_t is the real output and π^* and y^* are target levels of the inflation and output respectively. E is the expectation operator and Ω_t is the information available to the central bank at the time it sets the interest rate. In our paper we add the stock market into the equation above and get

$$R_t^* = \bar{R} + \beta(E[\pi_{t+n}|\Omega_t] - \pi^*) + \gamma(E[y_t|\Omega_t] - y_t^*) + \theta(E[A_{t+n}|\Omega_t] - A^*), \quad (28)$$

where A_{t+n} is the asset price in period $t+n$ and A^* denotes the fundamental value of the asset price. We expect θ to be positive, since we assume that central

banks try to stabilize the stock market with the interest rate as the instrument. If we define $\alpha = \bar{R} - \beta\pi^*$, $x_t = y_t - y^*$ and $B_{t+n} = A_{t+n} - A^*$ (namely the asset price bubble), equation (28) can be rewritten as

$$R_t^* = \alpha + \beta E[\pi_{t+n}|\Omega_t] + \gamma E[x_t|\Omega_t] + \theta E[B_{t+n}|\Omega_t], \quad (29)$$

after substituting equation (29) into (27), we have the following path for R_t :

$$R_t = (1 - \rho)\alpha + (1 - \rho)\beta E[\pi_{t+n}|\Omega_t] + (1 - \rho)\gamma E[x_t|\Omega_t] + (1 - \rho)\theta E[B_{t+n}|\Omega_t] + \rho R_{t-1} + v_t. \quad (30)$$

After eliminating the unobserved forecast variables from the expression, we get the following presentation:

$$R_t = (1 - \rho)\alpha + (1 - \rho)\beta\pi_{t+n} + (1 - \rho)\gamma x_t + (1 - \rho)\theta B_{t+n} + \rho R_{t-1} + \eta_t, \quad (31)$$

where $\eta_t = -(1 - \rho)\{\beta(\pi_{t+n} - E[\pi_{t+n}|\Omega_t]) + \gamma(x_t - E[x_t|\Omega_t]) + \theta(B_{t+n} - E[B_{t+n}|\Omega_t])\} + v_t$ is a linear combination of the forecast errors of the inflation, output gap, asset price bubble and the i.i.d. v_t . Let μ_t be a vector of variables within the central bank's information set at the time it chooses the interest rate that are orthogonal to η_t , we have

$$E[R_t - (1 - \rho)\alpha - (1 - \rho)\beta\pi_{t+n} - (1 - \rho)\gamma x_t - (1 - \rho)\theta B_{t+n} - \rho R_{t-1} | \mu_t] = 0. \quad (32)$$

GMM will be applied to estimate this equation with quarterly data of the EURO-area.¹⁷ Let $\pi_{t+n} = \pi_{t+4}$, as for B_{t+n} we will try with different n (0,1,..,4).¹⁸ The estimates with different n of B_{t+n} are presented in Table 3, with t-statistics in parentheses.

As is shown in Table 3, β and γ always have the right signs and significant t-statistics no matter which forward value of B_t is taken, which indicates that the inflation and output always play important roles in the interest rate setting. As for θ we find that it always has the right sign as expected, but the t-statistics is not always significant enough. When $n=0,1$ it is insignificant, when $n=3,4$ it is not very significant, but when $n=2$ it is significant enough. Therefore we may say that the asset price may have played some role (maybe not an important role) in the interest rate setting. The simulated interest rate with $B_{t+n} = B_{t+2}$ is presented together with actual interest rate in Figure 29.

¹⁷As for the details of GMM, the reader is referred to Hamilton (1994). In order to get the initial estimates of the parameters, we estimate the equation with traditional non-linear 2SLS methods first, since η_t is correlated to the independent variables. The instruments include the 1-4 lags of the output gap, inflation rate, German call money rate, asset price bubbles, nominal USD/ECU exchange rate and a constant, the data cover 1978.1-1984. The instruments are pre-whitened before the estimation. Data source: International Statistics Yearbook 2000 (online).

¹⁸Correction for MA(4) autocorrelation is undertaken, and j-statistics are also presented to see the validity of the over-identifying restrictions.

Parameter	Estimates				
	n=0	n=1	n=2	n=3	n=4
ρ	0.813 (19.792)	0.811 (18.561)	0.894 (30.224)	0.833 (15.870)	0.832 (17.089)
α	0.030 (4.581)	0.028 (3.920)	0.007 (0.466)	0.020 (1.918)	0.021 (2.074)
β	0.748 (5.446)	0.777 (5.343)	1.522 (3.921)	0.940 (4.410)	0.890 (4.567)
γ	2.046 (5.679)	2.011 (5.300)	1.626 (3.234)	2.345 (3.990)	2.363 (4.203)
θ	0.014 (0.509)	0.030 (0.927)	0.240 (2.328)	0.081 (1.264)	0.082 (1.100)
R^2	0.914	0.913	0.930	0.904	0.904
$J - stat.$	0.088	0.087	0.111	0.069	0.074

Table 3: GMM Estimates with Different n for B_{t+n}

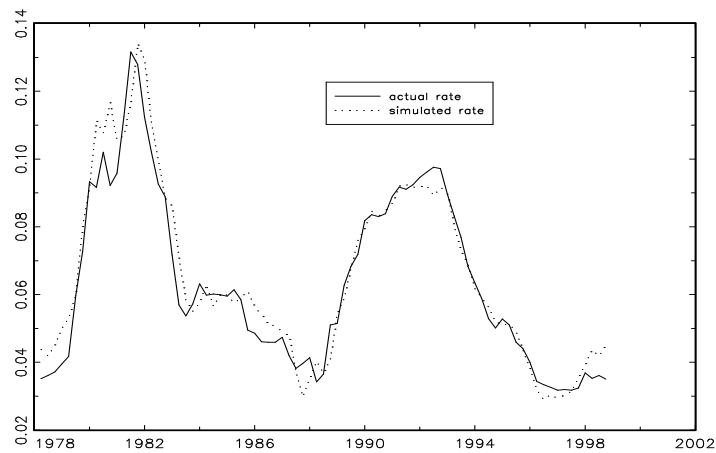


Figure 29: Actual and Simulated Interest Rates of the EURO-Area with Asset Price Bubble

5.3 Monetary Policy with Asset Price Targeting

From the estimation of the interest rate response equation with expectations we find that the stock market may have played a role in the interest rate determination in the EURO-area. Considering also the tests by Smets (1997) and Bernanke and Gertler (2000) for Australia, Canada, the U.S. and Japan, we tend to conclude that some countries seem to have taken into account the financial market volatility in monetary policy making in certain periods. It is, therefore, likely that the central banks may have included the asset price volatility as a direct target in their objective function. In Semmler and Zhang (2002) a dynamic model is set up to consider the monetary policy with and without the asset price targeting in the central bank's objective function. The simulations are undertaken for the U.S. and it is found that when asset price targeting is included into the objective function, the welfare loss is slightly decreased and the state variables (the inflation deviation, output gap and asset price bubble) all converge to their zero equilibria as time tends to infinity, no matter whether the loss function is symmetric or asymmetric and whether the probability of the asset price bubble to break is fixed or state-dependent. In the paper by Semmler and Zhang (2002) similar results for EURO-area estimations and simulations are reported. We will not present those results in detail here. The reader can be referred to the paper mentioned above.

6 Conclusion

This paper explores mainly the interest rate rules of the EURO-area economy. First we estimate the time varying Phillips-Curve to consider the Lucas-critique and then derive the monetary reaction function from a dynamic model, and explore whether the reaction function changes with the economic environment. We do find that the monetary reaction function experiences changes. On the basis of the research in the previous two sections, we explore whether the EURO-area interest rate was too tight in the 1990s. We undertake the simulations of inflation rate and output gap for the EURO-area using the following several interest rate rules: Time varying interest rate rule, fixed coefficient interest rate rule and the two so-called suggested Taylor rules of the U.S.. In order to make a comparison with the simulations undertaken with the four U.S. interest rate rules, we also undertake simulations with the actual EURO-area interest rate. All of the simulations seem to favor the conclusion that if EURO-area had followed the interest rate rules of the U.S. in the 1990s, the output gap would not have experienced such a fall and the output would have been higher. Of course, many observers of the monetary policy of the EURO-area would argue that lowering the interest rate would not have been a feasible policy since this would have led to an accelerated depreciation of European currencies and later of the EURO. As, however, shown in Semmler (2001) the EURO-area has large net foreign assets and thus large foreign currency reserves, so that an accelerated depreciation would not have occurred. Moreover, as recently has been shown

by Corsetti and Pesenti (1999) the high value of the dollar is strongly positively correlated with the growth differentials of the U.S. and Euro-area economies. A lower interest rate and thus higher expected growth rate of the EURO-area would have attracted capital inflows into the EURO-area and prevented the EURO from being depreciated.

The last section of the paper considers some issues on the relationship between the monetary policy and asset price volatility. We study whether central banks have and should take into account asset prices in their interest rate reaction function. The result favors the conclusion that the asset price volatility has been taken into account in the monetary policy determination. The welfare loss is, as Semmler and Zhang (2002) show, likely to be smaller as compared to a policy that neglects asset price bubbles.

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