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ABSTRACT

Asymmetries In Monetary Policy Rules: Evidence For Four Central Banks*

This Paper investigates the possible existence of asymmetric effects in the response of four central banks to inflation and output gaps as regards the 'sign' and 'size' of those gaps. The evidence obtained both through the estimation of a generalized Taylor rule and an ordered probit model points out that most central banks show a stronger reaction to inflation upswings relative to downswings. However, except for the Federal Reserve, no asymmetric behaviour with respect to the output gap is found.

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NON-TECHNICAL SUMMARY

This Paper studies the possible asymmetric nature of the preferences of central banks regarding inflation and output. The motivation of our work stems from some recent literature that suggests that the reaction of central banks, in terms of interventions through changes in a short-term interest rate, to deviations in the paths of both variables from their target levels may be asymmetric.

On the one hand, some authors have suggested that since central banks often deny the possibility that an expansionary policy may even have short-run real effects, say, on cyclical unemployment, it is likely that they will react differently with respect to positive and negative deviations of inflation from a target rate. On the other hand, it has been argued that central banks may have greater aversion to recession than to expansions and, therefore, may react with more virulence to negative output gaps than to positive ones.

To test for asymmetric behaviour, in the sense defined above, we use two empirical strategies. First, we estimate forward-looking Taylor rules allowing for different responses to both inflation and output gaps, according to their 'sign' and 'size'. Secondly, the above analysis is complemented with the estimation of an ordered probit model which analyses the determinants of the probabilities associated to the decisions on increasing, decreasing or keeping unchanged the interest rate under control. For both exercises, we use data from the European central banks (Bundesbank, Banque de France and Banco de España and from the Federal Reserve. The sample periods are 1980(8)–1997(12), 1988(7)–1997(12), 1989(5)–1997(12) and 1979(1)–1997(12), respectively.

As regards the inflation target, we depart from the standard assumption of constancy, and instead assume that it has time variation. The targets are obtained from various sources comprising: the Bundesbank reports for Germany; the German interest rate for France; the public budget laws for Spain; and the Council of Economic Advisors reports for the US. As for the output gaps, we use deviations from the Industrial Production Index in each country with respect to a HP filter.

The overall evidence seems to clearly support the belief that central banks intervene much more strongly when inflation moves above target than when it does below it. This has been particularly the case of the Bundesbank where an excess inflation of one percentage point above target triggers an increase of 1.8 percentage points in the interest rate, whereas there is hardly any reaction when the opposite case takes place. The results for the remaining central banks go in the same direction though they are weaker.

By contrast, with the exception of the Federal Reserve, which seems to react strongly to recessions, the three European central banks do not seem to react differently to ups and downs in the output gap.

1 Introduction

This paper is motivated by the work of Mishkin and Posen (1997) and Clarida and Gertler (1997) about the asymmetric nature of the preferences of central banks, which underlie the derivation of policy rules mimicking the conduct of monetary policy since the end of the 1970s. In contrast with the traditional result on the existence of an *´inflationary bias´* when central banks have discretionary power over the instrumentation of monetary policy, Mishkin and Posen suggest that a ´deflationary bias´ is a more likely outcome since independent central banks tend often to deny the possibility that an expansionary monetary policy stance may reduce cyclical unemployment. In particular, those authors argue in a descriptive manner that the behaviour of the Bank of Canada and the Bank of England is consistent with an asymmetric reaction with respect to positive and negative deviations of inflation from a target rate. Clarida and Gertler, in turn, test formally for the null hypothesis of symmetry and find evidence against it for the Bundesbank. Indeed, a careful reading of the definition of the target inflation rate in the euro-zone by the European Central Bank seems to point out towards an asymmetric stance, since such a target of price stability is defined as an HIPC inflation between 0 and 2%.

As for asymmetries regarding output, there is a new stream of literature pointing out that there is greater aversion to recessions than to expansions by central banks, as in Cuikerman (1999), or that the desired output level is a nonlinear function of shocks, as in Gerlach (2000). In both cases, the derived implication is an asymmetric response with respect to the output gap.

In this respect, the goal of this paper is to extend the previous evidence by using a generalised Taylor rule specification which encompasses the possible existence of asymmetric/non-proportional responses in the intervention behaviour of central banks in three European countries (France, Germany and Spain) and the US Federal Reserve.

By means of the previous framework, we develop simple tests for asymmetric behaviour with respect to both the inflation and output gaps, namely, the traditional explanatory variables in this type of policy rules; c.f. Taylor (1993), Clarida et al. (1997), Dornbusch et al (1998) and Peersman and Smets (1999) *inter alia.* Moreover, the analysis above is complemented by the estimation of an *ordered probit* model to analyse the probabilities associated to the decisions on increasing, decreasing or keeping unchanged a short-term nominal interest rate. In this way, we can assess through different econometric methods how different is the response of a central bank when there is an upsurge in inflation or output above their bliss points from the situation where there are negative gaps.

The rest of the paper is organised as follows. Section 2 reviews the derivation of forward looking monetary policy reaction functions both under symmetric and asymmetric preferences of central bank and discusses their estimable specifications. Section 3 presents the results, extending the analysis to allow for non-proportional response, namely, that a central bank may react differently to large and small changes in the determinants of monetary policy stance. Section 4 checks how robust are the previous results under a competing methodology based on an ordered probit model which studies the effects of positive and negative gaps on the probability of changing interest rate by different amounts. Finally, Section 5 concludes.

2 Derivation of the Taylor Rule

2.1 Symmetric loss function

In this section we lay out the basic principles which underlie the derivation of a linear ´forward looking´ Taylor rule following the arguments in Svensson (1997) and Clarida et al (1998, 1999). We will do that both for symmetric and asymmetric loss functions.

The central bank is assumed to control the monetary policy through the use of a policy rule. At time t, the monetary authorities commit to a state contingent sequence of short-term interest rates in order to minimize the following intertemporal loss function:

$$E_t \sum_{\tau=t}^{\infty} \delta^{\tau-T} L(\pi_\tau - \pi^*_\tau, y_\tau - y^*_\tau, i_\tau - \bar{i}) \tag{1}$$

s.t:

$$\pi_{t+k} = a_1 \pi_{t+k-1} + a_2(y_{t+p} - y^*_{t+p}) + u_{t+k} \tag{2}$$

$$(y_{t+p} - y^*_{t+p}) = b_1(y_{t+p} - y^*_{t+p})_{-1} - b_2(i_t - \pi_{t+1/t}) + \upsilon_{t+p+1} \tag{3}$$

where E_t is the conditional expectations operator, δ is the discount factor, $0<\delta < 1$, $k>p$ and $L(.)$ is a quadratic loss function given by:

$$L(.) = \lambda_1(\pi_\tau - \pi^*_\tau)^2 + \lambda_2(y_\tau - y^*_\tau)^2 + \lambda_3(i_\tau - \bar{i})^2 \tag{4}$$

such that $(\pi_\tau - \pi^*_\tau)$, $(y_\tau - y^*_\tau)$, represent deviations of the inflation rate (π) and the output gap (y_τ) with respect to a target inflation rate (π^*) and potential (y^*), respectively. Moreover, the central bank tries to approach a long-run equilibrium nominal rate, $\bar{i}$. As regards equation (2) it can be interpreted in terms of a Phillips curve where inflation is sluggish and depends on the cyclical component of output, whilst (3) can be thought of as an aggregate demand curve where output depends of its lagged value and on the real interest rate. Note that the nominal interest rate affects output with p-period lag and therefore affects inflation with a k-period lag.

As Svensson (1997) has proved, since i_t does not affect y_t and π_t contemporaneously, minimizing (1) turns out to be equivalent to the period-by-period static minimization:

$$\min_{i_t} E_t(\lambda_1(\pi_{t+k} - \pi^*_{t+k})^2 + \lambda_2(y_{t+p} - y^*_{t+p})^2 + \lambda_3(i_t - \bar{i})^2) \qquad (5)$$

Thus, the first order necessary conditions yield the following policy reaction function:

$$i^*_t = \bar{i} + \beta_1(E_t(\pi_{t+k}) - \pi^*_{t+k}) + \beta_2 E_t(y_{t+p} - y^*_{t+p}) \qquad (6)$$

Lastly, it is assumed that the central bank smooths interest rate changes by adopting an AR(1) partial adjustment rule such that[1]:

$$i_t = \rho i_{t-1} + (1-\rho) i^*_t \qquad (7)$$

[1] For the Federal Reserve, an AR(2) model was needed to pass the test for overidentifying restrictions.

Thus, the coefficient in the Taylor rule are determined by the underlying parameters governing the dynamics of inflation and output (a_1, a_2, b_1, b_2), the weights in the loss functions (λ_i, i=1...4) and the smoothing parameter (ρ). Accordingly, the $\beta_i's$ coefficients are expected to be positive and, in particular, if $\beta_1 > 1$ the target real rate $(i_t^* - E_t\pi_{t+k})$ adjust to stabilize inflation, as well as output (if $\beta_2 > 0$) whereas if $\beta_1 < 1$, instead it moves to accommodate inflation changes; c.f. Clarida et al (1997, 1999).

2.2 Asymmetric loss function

Once the standard linear Taylor rule has been derived, we turn to the case where the central bank does not weigh up equally positive and negative deviations of inflation from its target rate. In this case, the static loss function in (5) becomes:

$$\min_{i_t} E_t(\lambda_1^+[\max(\pi_{t+k} - \pi_{t+k}^*, 0)]^2 + \lambda_1^-[\min(\pi_{t+k} - \pi_{t+k}^*, 0)]^2 \quad (8)$$
$$+\lambda_2(y_{t+p} - y_{t+p}^*)^2 + \lambda_3(i_t - \bar{i})^2)$$

Proceeding in the same way as in the previous subsection leads to the following policy rule:

$$i_t^* = \bar{i} + \beta_1^+(E_t(\pi_{t+k}) - \pi_{t+k}^*)^+ + \beta_1^-(E_t(\pi_{t+k}) - \pi_{t+k}^*)^- + \beta_2 E_t(y_{t+p} - y_{t+p}^*) \quad (9)$$

Likewise, we can assume an asymmetric response with respect to the output deviations, which in this case yields the policy rule:

$$i_t^* = \bar{i} + \beta_1(E_t(\pi_{t+k}) - \pi_{t+k}^*) + \beta_2^+ E_t(y_{t+p} - y_{t+p}^*)^+ + \beta_2^- E_t(y_{t+p} - y_{t+p}^*)^- \quad (10)$$

2.3 Estimation of a Taylor Rule

According to the derivation in the previous section, the estimable specifications of the symmetric and asymmetric versions of the Taylor rule become:

$$i_t = \rho i_{t-1} + (1-\rho)\left\{\bar{i} + \beta_1(\pi_{t+k} - \pi_{t+k}^*) + \beta_2(y_{t+p} - y_{t+p}^*) + \beta_3' X_t\right\} + \varepsilon_t \quad (11)$$

and

$$i_t = \rho i_{t-1} + (1-\rho)\left\{\begin{array}{c} \bar{i} + \beta_1^+(\pi_{t+k} - \pi_{t+k}^*)^+ + \beta_1^-(\pi_{t+k} - \pi_{t+k}^*)^- + \\ \beta_2(y_{t+p} - y_{t+p}^*) + \beta_3' X_t \end{array}\right\} + \varepsilon_t \quad (12)$$

where X_t denotes a set of observable variables, besides inflation and output gaps, that may potentially affect interest rate setting independently of their role in helping to forecast the above-mentioned variables. They include, for instance, variations in real exchange rate, foreign interest rate and the evolution of the money supply. Notice that, by exploiting the rational hypothesis in expectations formation we can replace forecast variables by their realized values so that the error term follows the stochastic process:

$$\varepsilon_t = \xi_t - (1-\rho)(\beta_1 e_{t+k/t}^{\pi} + \beta_2 e_{t+p/t}^{y}) \quad (13)$$

where ξ_t is an $i.i.d$ disturbance, $e_{t+k/t}^{\pi} \equiv \pi_{t+k} - E(\pi_{t+k})$ is the k-period ahead forecast error for inflation and $e_{t+p/t}^{y} \equiv y_{t+p} - E(y_{t+p})$ is the corresponding p-period ahead forecast error for the output gap. Finally, let z_t be a vector of variables within the central bank´s information set, such as lagged variables that help forecast inflation and output or any other contemporaneous variables that are uncorrelated with the policy rule shock, ε_t. Then, the Generalized Method of Moments (GMM) can be used to estimate the parameter vector in (6), (7) and (11) by exploiting the set of orthogonality conditions.

$$E(\varepsilon_t / z_t) = 0 \tag{14}$$

Further, since the composite disturbance has an MA(*k-1*) representation due to the overlapping nature of the forecast errors, the weighting var-cov matrix used to implement GMM is the one proposed by Newey and West (1987). Moreover, Hansen (1982) ´s J test is used to test the overidentification restrictions and the null hypothesis $\beta_i^+ = \beta_i^-$ is tested to check for the existence of asymmetries.

3 Results

We estimate the monetary policy reaction functions for four central banks: the Bundesbank, the Bank of France and the Bank of Spain in Europe and the Federal Reserve in the US. The sample periods has been determined on the basis of choosing homogeneous spells where there was a virtually autonomous control over domestic monetary policy in each case. So, they correspond to 1980(8)-1997(12) for Germany, 1988(7)-1997(12) for France, 1989(5)-1997(12) for Spain and 1979(1)-1997(12) for US.

As for the short-term interest rates, they are chosen as follows: (i) the three-month interbank market rate in Germany and France, (ii) the marginal intervention rate of auctions of "Certificados del Banco de España" in Spain, and (iii) the Fed-Fund rate in the US. Inflation is measured through the CPI inflation rate and output through the (log of) Industrial Production Index. To obtain a measure of output gap, we detrend the log of industrial production index using the Hodrick-Prescott(HP) filter. As regards the inflation target, π^*, we depart from the assumption that it is constant, as in Clarida et al.(1997) and instead assume that it has time variation according to the following considerations: (i) in the case of Germany, we take the inflation target to be the one established by the Bundesbank in its annual reports; (ii) in the case of France, we take it to be the German target inflation rate, given the close links between both economies within the EU; (iii) in the case of Spain, it is proxied by the official inflation rate in the budget laws up to 1995 and the target inflation rate of the Bank of Spain since 1996; and (iv) in the case of US, we take it to be the one in the reports of the Council of Economic Advisors.

In all instances the annual target rates have been interpolated to a monthly frequency. Notice that the choice of a time-varying target rate is sensible for the analysis of asymmetries since some of the countries in the sample have experienced long disinflationary periods making it difficult to believe that a constant long-term inflation rate was guiding monetary policy in the short-run. Figure 1 plots the inflation rates and their assumed targets in each of the four countries.

3.1 Symmetric specification

Table 1 reports the results for the symmetric specification of the policy rule. We begin with the Bundesbank, where the list of instruments includes a constant term and six lags of the following variables: inflation, output gap and interest rate (in the first two columns) and two lags of the DM/$ real exchange rate (*rer*), US interest rate and deviations of M3 with respect to its target level, which are variables included in the X_t set. As for the choice of k and p we report two specifications, one with *k=12* and *p=0* and another with *k=12* and *p=6* (second column). Finally, the case where *k=12* and *p=0*, which turned out to have the best fit best the data, was augmented with the X_t set of variables as described above[2]. Having added each of the X_t variables one by one, we found that the specification containing the US interest rate was the only one together with the one containing the deviations of M3 with respect to its target level, albeit marginally so, where the coefficient was significant. The results in the three columns are fairly similar, pointing out that the degree of persistence is very high ($\rho \approx 0.9$) and that the Bundesbank responds to the inflation gap more than proportionally ($\beta_1 = 1.45$ in the baseline specification). The estimate of the output gap is positive and significant ($\beta_2 = 0.28$).

With regard to the Bank of France, the preferred specification was the

[2] The list of instruments for the equations pertaining to the Bank of France and the Bank of Spain ate the same as before except that the DM/$ exchange rate is replaced by the FF/DM and the Pta/DM exchange rate, respectively, and the foreign interest rate is the German one. As for the Federal Reserve, the additional variables are two lags of total and non-borrowed reserves.

one that contained the change of the FF/DM real exchange rate (rer) and the German interest rate (i_t^G) within the X_t set. Given the estimate of the latter variable ($\beta_{32} = 0.87$) we can interpret the policy rule as a weighted average of the German interest rate (0.87) and the baseline policy rule (0.13). One big difference with the Bundesbank is that the Bank of France has responded to the inflation gap less than proportionally ($\beta_1 = 0.41$), indicating the monetary policy stance has been fairly accommodating, with the increases in the short-term interest rates not being of enough size to keep the real interest rate from declining. Moreover the response to the output gap is much weaker than in Germany ($\beta_2 = 0.13$).

As regards the Bank of Spain, the results are fairly similar to the ones found for the French case with the response to the inflation gap being smaller than unity ($\beta_1 = 0.53$). However, the response to the output gap is about the same size that in Germany ($\beta_2 = 0.45$). Finally, the size of the estimated coefficient on the German interest rate is slightly larger than unity, a result which is difficult to explain within the framework of a linear combination between a policy rule and the short-term interest rate as in Germany, but which, however, indicates that the Bank of Spain has been very closely shadowing the German interest rate since acceeding to the ERM in 1986.

Finally, the results for the Federal Reserve are similar to those obtained for the German policy rule ($\beta_1 = 1.21$), albeit the response to the output gap is much stronger than in any of the European countries. Within the X_t variables, both contemporaneous changes in total reserves (tr) and in non-borrowed reserves (nbr) turned out to be statistically very significant, triggering increases in the Fed-Fund rate.

In most of the specifications the J-test takes p-values above 0.05, non rejecting the set of overidentifying restrictions.

3.2 Asymmetric specification

Table 2 reports the results for the asymmetric specification concerning deviations of future inflation with respect to the target rate. As before, for each of the central banks we report three alternative models corresponding to $p=0$ or $p=6$ and the inclusion of X_t variables. The reported results for the Bundesbank show very clearly that the response of interest rates to inflation deviations above their target is much stronger than when it is below target, in which case no estimates are statistically significant. Thus, the null hypothesis $\mathrm{H}_0 : \beta_1^+ = \beta_1^-$ turns out to be rejected at very low p-values, as shown in the fourth column of the panel. The size of the estimated coefficients points out to a clear policy of raising the real interest rate, by about 80 b.p. when inflation is 1 p.p. above its target, while allowing for a real interest increase of about 70 b.p., when it is 1 p.p below it.

As regards the Bank of France, we obtain that it responds to both types of deviations although again more strongly to positive deviations. In both cases, however, in agreement with the results in Table 1, the policy rule turns out to destabilise inflation. The null hypothesis of symmetry is again rejected.

In the case of the Bank of Spain, due to the low number of observations where inflation has been below its target rate, we have followed a slightly different route to that specified in equation (9). Indeed, we have replaced the inflation gap by the change in inflation, distinguishing between price level accelerations and decelerations, so that (9) becomes:

$$i_t = \rho i_{t-1} + (1-\rho)\left\{ \begin{array}{c} \bar{i} + \beta_1^{+}(\Delta\pi_{t+k})^{+} + \beta_1^{-}(\Delta\pi_{t+k})^{-} + \\ \beta_2(y_{t+p} - y_{t+p}^{*}) + \beta_3' X_t \end{array} \right\} + \varepsilon_t \qquad (15)$$

The null hypothesis is rejected both in the first and third specifications. Somewhat surprisingly we find that $\beta_1^{+} < \beta_1^{-}$, a result which is difficult to explain for a country whose goal has been to disinflate over the sample period.

With regard to the Federal Reserve, we find a similar pattern of responses to that found for the Bundesbank, though β_1^{-} turns out to be statistically significant.

Figure 2 portrays the fitted values of the symmetric and asymmetric model[3]. It thus appears that the asymmetric model predicts the evolution of the short-term interest rate better than the symmetric model, particularly in the cases of France and Spain.

Finally, though not reported for the sake of brevity, we tested for symmetric responses with respect to positive and negative output gaps. In none of the three European central banks we found any evidence against symmetry. However, for the Federal Reserve we found that while β_2^{+} (0.07) was statistically insignificant, while β_2^{-} (1.75) was significant, rejecting the null hypothesis with a p-value of 0.03. This evidence is in agreement with the results in Gerlach (2000). Yet, using a slightly different specification of the Taylor rule, he only obtains that the Federal Reserve responds more strongly to negative output gaps during the period 1960-1979.

[3] We chose for each case the specification with the highest p-value.

3.3 Policy rules with non-proportional responses

In this section we turn to test whether there are significant differences in the way central banks respond to the size of the inflation gap. This type of hypothesis underlies the common belief that monetary authorities tend to have a stronger policy stance when faced with large deviations from inflation and output from their target rates than when those are small.

To undertake this test, we modify the baseline specification of the policy rule to allow for a non-proportional response in the following form:

$$i_t = \rho i_{t-1} + (1-\rho) \left\{ \begin{array}{c} \bar{i} + (\theta_0 + \theta_1 \left| \pi_{t+k} - \pi^*_{t+k} \right|)(\pi_{t+k} - \pi^*_{t+k}) + \\ \beta_2 (y_{t+p} - y^*_{t+p}) + \beta_3' x_t \end{array} \right\} + \varepsilon_t \quad (16)$$

where $|.|$ stands for the absolute value of the inflation gap. The null hypothesis of proportional response corresponds to $\mathrm{H}_0 : \theta_1 = 0$. If H_0 is rejected and $\theta_1 > 0$ we would interpret the outcome of the test as a ´more than proportional´ response and conversely for $\theta_1 < 0$.

The results for this test are shown in Table 3. In every case the null hypothesis was rejected, though in Spain only at the 6% level. For the three European countries the evidence is favorable to a ´more than proportional´ response whereas the opposite seems to be the case in the US.

As for the possibility of having non-proportional responses with respect to the output gap, we do find again no evidence in its favour.

4 Ordered probit models

In this section we concentrate on checking the robustness of the results obtained in the previous sections to the use of a non-linear approach based upon the estimation of an ordered probit model to analyse the determinants of interest rate changes from a slightly different perspective. Through such a methodology, we assume that the monetary authority takes a decision every month about implementing the following interventions in terms of interest rate changes: large or small reductions, keeping them invariant, and large and small increases.

4.1 Model analysis

As is well known, and ordered probit generalises the linear regression model to the case of a discrete choice dependent variable. Accordingly, we will assume that changes in interest rates are discrete and use a breakdown of interventions in the following five categories:

$$
\begin{aligned}
\text{(large decrease) } c_t &= 1 \iff \Delta i_t < -0.25 \\
\text{(small decrease) } c_t &= 2 \iff \Delta i_t \in [-0.25, 0) \\
\text{(inactivity) } c_t &= 3 \iff \Delta i_t = 0 \\
\text{(small increase) } c_t &= 4 \iff \Delta i_t \in (0, 0.25] \\
\text{(large increase) } c_t &= 5 \iff \Delta i_t > 0.25
\end{aligned}
$$

The constructed dummy variable, c_t, depends on a latent variable, c_t^*, according to the following rule:

$$c_t = \begin{cases} 1 \text{ if } c_t^* \leq \alpha_1 \\ 2 \text{ if } \alpha_1 < c_t^* \leq \alpha_2 \\ 3 \text{ if } \alpha_2 < c_t^* \leq \alpha_3 \\ 4 \text{ if } \alpha_3 < c_t^* \leq \alpha_4 \\ 5 \text{ if } c_t^* > \alpha_4 \end{cases}$$

where c_t^* is taken to be a continuous random variable which depends linearly on a set of covariates, x_t, such that

$$c_t^* = \beta' x_t + \varepsilon_t \tag{17}$$

with ε_t following a n.i.d. $(0,\ \sigma_\varepsilon^2)$. Such an assumption leads to an ordered probit model (see Maddala, 1983). From the previous assumptions, the probabilities of observing a given value of c_t are given by:

$$\begin{aligned} p(c_t &= 1/\alpha, \beta, x) = \Phi(\alpha_1 - \beta' x_t) \qquad (18) \\ p(c_t &= j/\alpha, \beta, x) = \Phi(\alpha_j - \beta' x_t) - \Phi(\alpha_{j-1} - \beta' x_t)\ \forall j = 2, 3, 4 \\ p(c_t &= 5/\alpha, \beta, x) = 1 - \Phi(\alpha_4 - \beta' x_t) \end{aligned}$$

where $\Phi(.)$ is the cumulative gaussian distribution function. Estimates of the parameter vector $(\alpha,\ \beta)$ are then obtained through maximization of the following likelihood function:

$$\begin{aligned} l(\alpha, \beta) \quad = \quad & \sum_{t \in y_t = 1} \log \Phi(\alpha_1 - \beta' x_t) + \sum_{j=2}^{4} \sum_{t \in y_t = j} \log(\Phi(\alpha_j - \beta' x_t) - \Phi(\alpha_{j-1} - \beta' x_t)) + \\ & \sum_{t \in y_t = 5} \log(1 - \Phi(\alpha_4 - \beta' x_t)) \end{aligned}$$

4.2 Results

The underlying response model used in the estimation is:

$$
\begin{aligned}
c_t^* &= \beta' x_t + \varepsilon_t = \beta_1(E_t(\pi_{t+k}) - \pi_{t+k}^*) + \beta_2(y_{t+p} - y_{t+p}^*) + \beta_3 \Delta rer_t \quad (19) \\
&\quad + \beta_4 \Delta i_{t-1} + \beta_5 \Delta i_{t-1}^{ge} + \beta_6 D_t + \varepsilon_t
\end{aligned}
$$

where Δi_{t-1} represents the interest rate change in the previous month and D_t denotes the number of months elapsed since the last intervention; c.f. Dolado and María-Dolores (2000). The specification for the Bundesbank includes changes in the DM/$ lagged real exchange rate (rer), whereas the one for central banks of France and Spain contains changes in the German interest rate (i^G); finally, the equation for the Federal Reserve contains changes in total and non-borrowed reserves. Since the regressors in the probit model are assumed to be uncorrelated with the error term, the procedure of replacing expectations by their realised value becomes invalid. Thus, rather than using the previous approach, our strategy is based on constructing inflation and output forecasts from OLS regression equations where the regressors are the instrumental variables used in the GMM approach, namely, lags on inflation, output gap, changes of real exchange rate, raw materials price index and deviations of money growth with respect to a target rate.

Table 4 shows the results of the above exercise. We find that the probability of intervention increases with the inflation and the output gap (except in France), when the real exchange rate depreciates, when the duration since the last intervention increases and when the lagged interest rate increased. Next, we allow for the different responses to vary depending on whether inflation is above or below its target rate and compute the marginal effects of positive and negative deviations of inflation on the five different probabilities of intervening, according to the following expressions:

$$\frac{\partial prob(y_t = 1)}{\partial X_i} = \phi(\alpha_1 - \beta X_{it})(-\beta_i) \qquad \forall i = 1...k \tag{20}$$

$$\frac{\partial prob(y_t = 5)}{\partial X_i} = \phi(\alpha_5 - \beta X_{it})(-\beta_i) \qquad \forall i = 1...k \tag{21}$$

$$\frac{\partial prob(y_t = j)}{\partial X_i} = [\phi(\alpha_k - \beta X_{it}) - \phi(\alpha_{k-1} - \beta X_{it})](-\beta_i) \qquad \forall i = 1...k \quad \forall j = 2, 3, 4 \tag{22}$$

$$\frac{\partial prob(y_t = k)}{\partial(\hat{\pi}_{t+k/t} - \pi^*_{t+k})^+} = \left[\phi(\alpha_k - \beta\bar{x}^+) - \phi(\alpha_{k-1} - \beta\bar{x}^+)\right](-\beta_1^+) \tag{23}$$

$$\frac{\partial prob(y_t = k)}{\partial(\hat{\pi}_{t+k/t} - \pi^*_{t+k})^-} = \left[\phi(\alpha_k - \beta\bar{x}^-) - \phi(\alpha_{k-1} - \beta\bar{x}^-)\right](-\beta_1^-) \tag{24}$$

where:

$\bar{x}^+ =$ sample average of observations with $(\hat{\pi}_{t+k/t} - \pi^*_{t+k})^+ \geqslant 0$

$\bar{x}^- =$ sample average of observations with $(\hat{\pi}_{t+k/t} - \pi^*_{t+k})^- < 0$

The estimated marginal effects are presented in Table 5. We basically replicate the results obtained with the Taylor rule approach with a few exceptions. Thus, for instance, we get that a positive 1 p.p deviation of inflation from target increases the probability of raising interest rates by more than 25 b.p. in 0.08. By contrast, when inflation is 1 p.p. below target, the probability of decreasing interest rate by more than 25 b.p. only falls by 0.004.

Similar results are obtained for the three remaining countries, where the difference in probabilities is decreasing as the perceived change in interest rates is lower. In this respect, the new evidence for the Bank of Spain contradicts the previous one where, according to the Taylor rule approach, it seemed that the central bank was more active when facing negative inflation deviations than the opposite. Thus, whereas the evidence for Spain looks inconclusive, the overall results for the other three countries confirm the presence of important asymmetries in the behaviour of central banks leading to a ´deflationary bias´, i.e. taking an active contractionary stance when the economy suffers an ´overheating´ but remaining less active when it ´cools off´.

Finally, through not reported, we also implemented a similar exercise looking for non-proportional responses. Once again, we hardly found any evidence in its favour, confirming the earlier evidence.

5 Concluding remarks

In this paper we searched for asymmetric responses to inflation and output gaps in the policy responses of four central banks. For that purpose, we have developed two statistical testing procedures. The first one is based on a generalised Taylor rule which allows for asymmetric effects of both positive and negative gaps when determining interventions in terms of changes of a short-run nominal interest rate. The second one is based on an ordered probit model which captures the discrete nature of those changes and therefore models the probability of implementing a series of different interventions as a function of the perceived state of the economy.

The overall evidence seems to clearly support the belief that central banks intervene with much more virulence when inflation moves above its target than when it does below it. This has been particularly the case of the Bundesbank where an excess inflation of 1 p.p. above target triggers an increase of 1.8 p.p in the interest rate whereas there is hardly any reaction when the opposite case takes place. The results for the central banks of France, Spain and the US go in the same direction though they are weaker.

By contrast, with the exception of the Federal Reserve, which seems to react strongly to recessions, the three European central banks do not react differently to ups and downs in the output gap.

In sum, the above evidence seem to confirm the hypothesis posed by Mishkin and Posen (1997) and Clarida and Gertler (1997) that there might be a ´deflationary bias´ in the operating procedures of central banks. Moreover, the statistically significant difference in the policy rules may turn out to be helpful for financial market analysts when forecasting future changes in monetary stance on the basis of the already very popular usage of Taylor rules.

Table 1: Estimated Reaction functions for Central Banks

	GE			FR			SP			US		
	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12	k = 12
	p = 0	p = 6	p = 6	p = 0	p = 6	p = 6	p = 0	p = 6	p = 6	p = 0	p = 6	p = 6
½$_1$	:94	:96	:95	:85	:86	:87	:93	:94	:96	1:25	1:24	1:17
	(:02)	(:01)	(:02)	(:02)	(:02)	(:02)	(:02)	(:02)	(:02)	(:07)	(:06)	(:06)
½$_2$	-	-	-	-	-	-	-	-	-	-0.32	-0.31	-0.37
	-	-	-	-	-	-	-	-	-	(.06)	(.07)	(.06)
ł	5:75	5:77	5:84	5:72	5:68	5:81	5:86	5:93	6:05	7:20	7:21	6:94
	(:44)	(:53)	(:49)	(:29)	(:33)	(:34)	(:84)	(:64)	(:98)	(:73)	(:82)	(:38)
—$_1$	1:45	1:31	1:09	:41	:39	:42	:53	:49	:51	1:21	1:23	1:81
	(:63)	(:65)	(:31)	(:14)	(:15)	(:11)	(:20)	(:18)	(:15)	(:32)	(:35)	(:14)
—$_2$	:28	:47	:54	:13	:09	:21	:45	:20	:45	:90	:79	:82
	(:13)	(:23)	(:21)	(:07)	(:06)	(:17)	(:33)	(:11)	(:26)	(:41)	(:31)	(:24)
—$_{31}$	¡	¡	¡	:19	:23	:25	:91	:93	:92	¡	¡	:05
	¡	¡	¡	(:11)	(:14)	(:15)	(:57)	(:54)	(:55)	¡	¡	(:006)
—$_{32}$	¡	¡	:14	:87	:82	:85	1:23	1:19	1:52	¡	¡	:06
	¡	¡	(:06)	(:19)	(:21)	(:22)	(:40)	(:31)	(:71)	¡	¡	(:006)
¾"	.34	.35	.31	.69	.71	.70	.71	.73	.72	.78	.80	.79
p	.09	.04	.08	.11	.16	.14	.13	.04	.16	.16	.07	.11

GE~Germany $X_t = i_t^{USA}$ Sample Period 1980:08 1997:12

FR~France X_t = ¢rer$_t$; i_t^G Sample Period 1988:07 1997:12

SP~Spain X_t = ¢rer$_t$; i_t^G Sample Period 1989:05 1997:12

US~USA X_t = ¢tr$_t$; ¢nbr$_t$ Sample Period 1979:01 1997:12

Note: Standard errors in parentheses. ½; is the p-value of the J-test for overidentifying restrictions

Table 2: Test for asymmetric response to inflation gap

	Germany			France			Spain			US		
$H_o : \beta_1^+ = \beta_1^-$	β_1^+	β_1^-	p	β_1^+	β_1^-	p	β_1^+	β_1^-	p	β_1^+	β_1^-	p
$k = 12$	1.86	.26	.007	.65	.23	.04	.10	.39	.04	1.11	0.80	.002
$p = 0$	(.32)	(.20)		(.16)	(.07)		(.04)	(.10)		(.45)	(.35)	
$k = 12$	1.79	.19	.006	.68	.21	.04	.09	.32	.07	1.06	0.60	.001
$p = 6$	(.35)	(.16)		(.15)	(.09)		(.04)	(.11)		(.48)	(.29)	
$k = 12$, $p = 6$, $+ X_t$	1.68	.29	.008	.62	.19	.03	.08	.37	.05	1.13	0.57	.002
	(.28)	(.22)		(.13)	(.08)		(.04)	(.12)		(.23)	(.23)	

Note: Standard error in parentheses; p is the value of $H_0 : \beta_1^+ = \beta_1^-$

Table 3: Test for non-proportional response to inflation gap			
$H_o : \theta_1 = 0$	θ_0	θ_1	p
Germany	.87 (.37)	.10 (.05)	.039
France	.39 (.15)	.11 (.04)	.018
Spain	.44 (.19)	.08 (.04)	.053
USA	3.09 (.74)	$-.16$ (.01)	.005
Note: Standard error in parentheses;p is the value of $H_0 : \theta_1 = 0$			

Table 4: Ordered Probit Model

Variables	Germany	France	Spain[a]	USA
$\hat{\pi}_{t+12/t} - \pi^*_{t+12}$	.13	.21	.43	0.11
	(.06)	(.09)	(.15)	(.06)
$\hat{y}_{t+6/t} - y^*_{t+6}$	.08	$-.01$	.08	.11
	(.02)	(.04)	(.03)	(.04)
Δrer_t	.14	.78	.04	–
	(.03)	(.32)	(.07)	–
Δi^G_{t-1}	–	2.29	.60	–
	–	(.53)	(.27)	–
Δi_{t-1}	.81	.57	0.69	.24
	(.21)	(.25)	(.21)	(.12)
D_t	.18	.14	.13	.38
	(.12)	(.06)	(.06)	(.30)
Δtr_t	–	–	–	.04
	–	–	–	(.001)
Δnbr_t	–	–	–	.06
	–	–	–	(.005)
α_1	-1.12	$-.92$	-1.18	$-.78$
	(.10)	(.15)	(.18)	(.10)
α_2	$-.33$	$-.07$	.18	$-.12$
	(.09)	(.04)	(.10)	(.09)
α_3	.31	.42	1.33	.19
	(.09)	(.13)	(.18)	(.09)
α_4	1.07	1.36	1.84	.83
	(.10)	(.17)	(.22)	(.10)
$Log - Likelihood$	-334.576	-155.498	-124.447	-339.452

[a]The variable $\pi_{t+12/t} - \pi^*_{t+12}$ is replaced by $\Delta\pi_{t+12/t}$

Note: Standard errors in parentheses

Table 5: Asymmetric responses to inflation in the ordered probit model

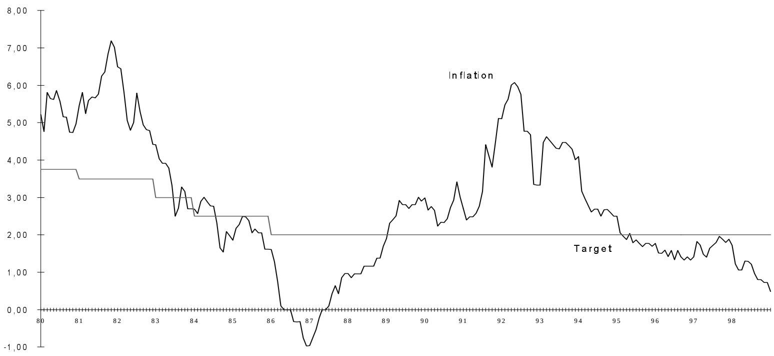
	$\frac{\partial prob(c_t=1)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^-}$	$\frac{\partial prob(c_t=5)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^+}$	$\frac{\partial prob(c_t=2)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^-}$	$\frac{\partial prob(c_t=4)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^+}$	$\frac{\partial prob(c_t=3)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^-}$	$\frac{\partial prob(c_t=3)}{\partial(\hat{\pi}_{t+k/t}-\pi^*_{t+k})^+}$
Germany	.0043	.081	.011	.076	.0037	.014
France	.011	.078	.024	.066	.0064	.032
Spain	.037	.074	.065	.096	.012	.021
US	.0017	.022	.0023	.024	.0025	.022

[a]In Spain the variable $(\hat{\pi}_{t+12/t} - \pi^*_{t+12})$ has been replaced by $\Delta\hat{\pi}_{t+12/t}$

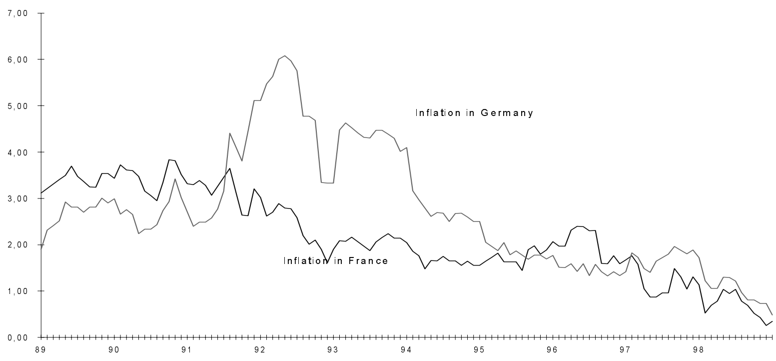
Note: The derivations are evaluated in the sample means of the explanatory variable

Figure 1

Current and Target Inflation Rates

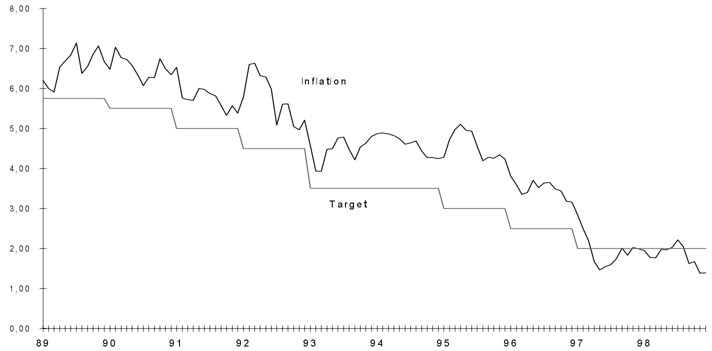


Inflation and Target in Germany

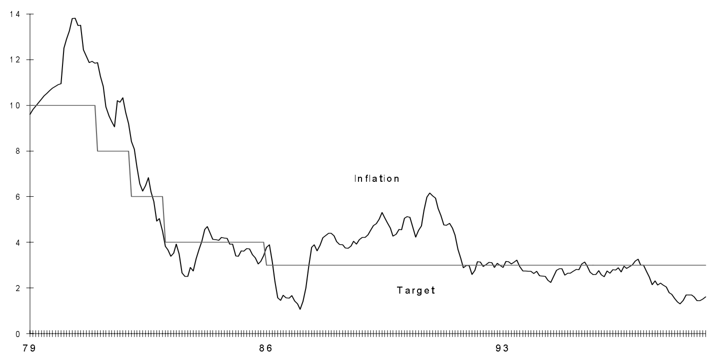


Inflation and Target in France

Figure 1 (Cont.)



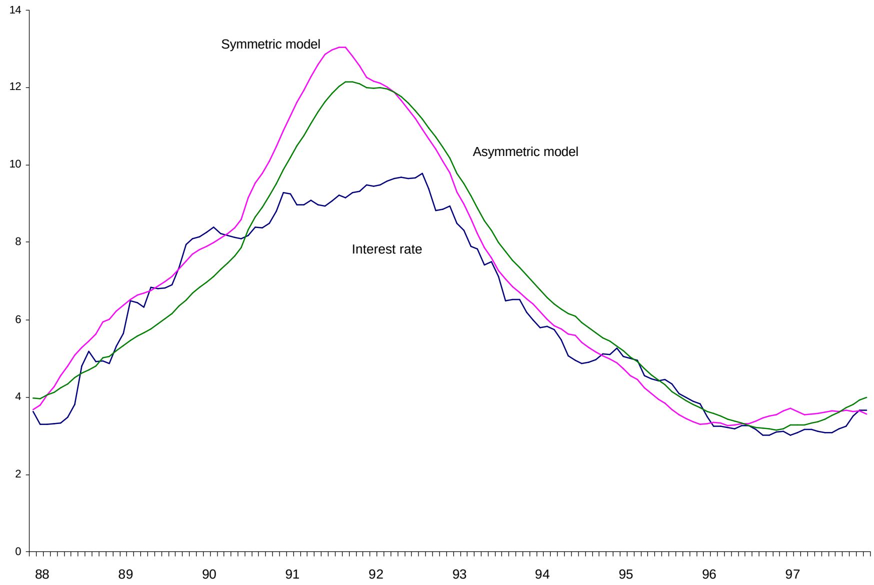
Inflation and Target in Spain



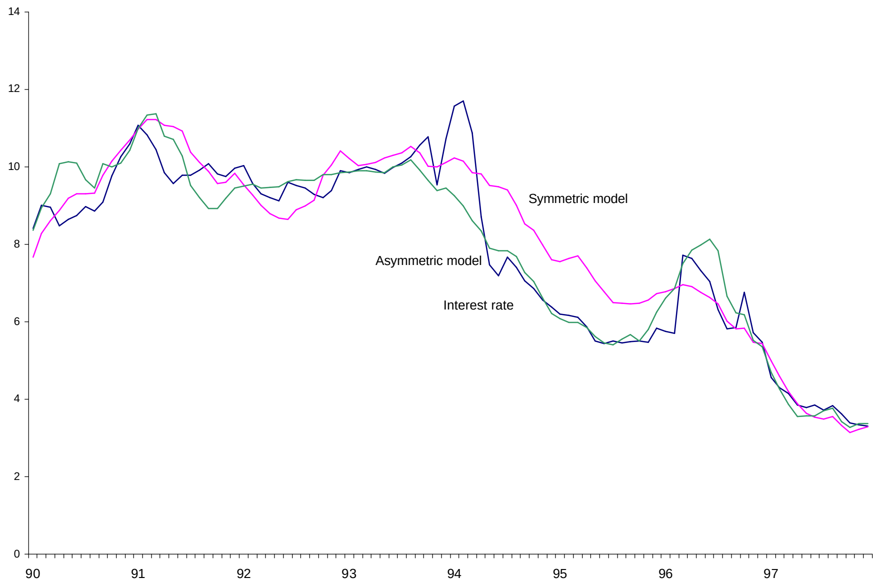
Inflation and Target in US

Figure 2

Asymmetric and symmetric model predictions



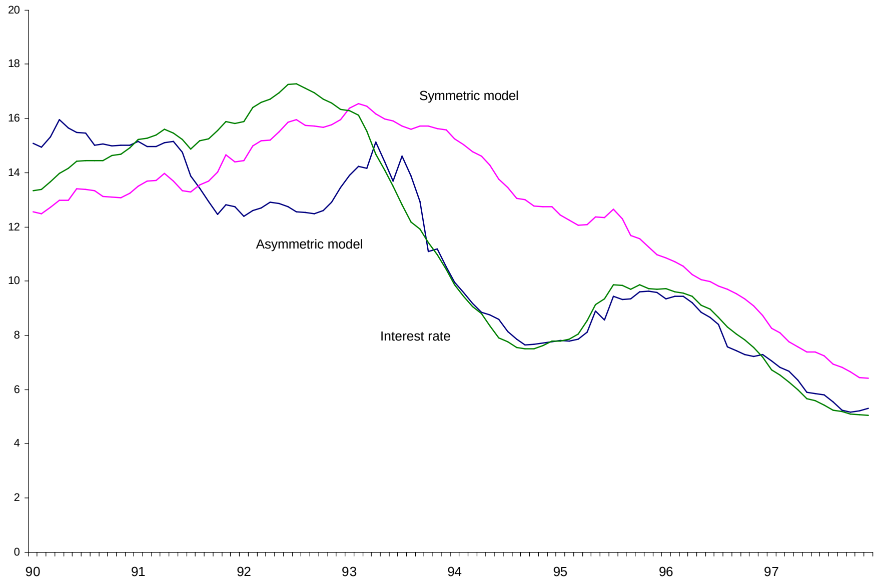
Taylor rule predictions in Germany



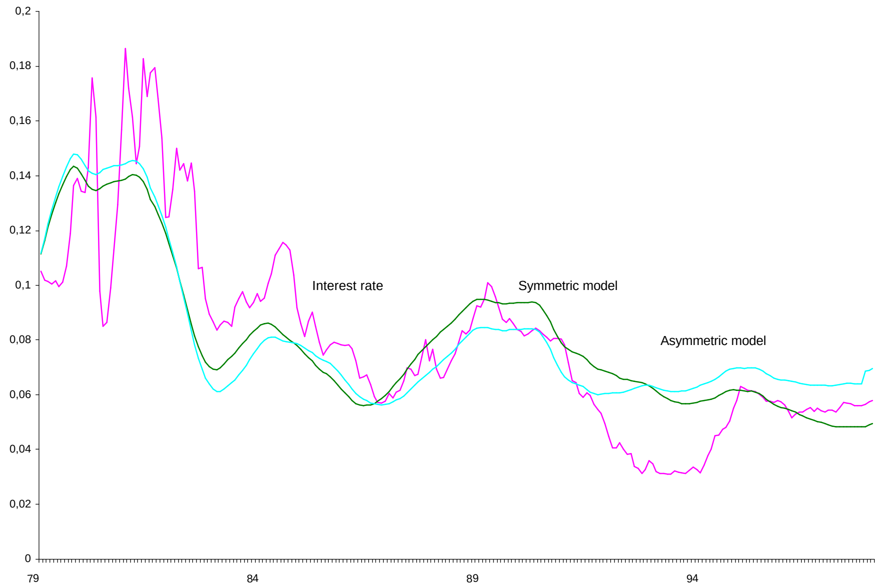
Taylor rule predictions in France

Figure 2 (Cont)

Asymmetric and symmetric model predictions



Taylor rule predictions in Spain



Taylor rule predictions in USA

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