

Zentrum für Europäische Integrationsforschung  
Center for European Integration Studies  
Rheinische Friedrich-Wilhelms-Universität Bonn



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**MONETARY POLICY  
REACTION FUNCTIONS:  
ECB VERSUS BUNDESBANK**

Working Paper

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# Monetary Policy Reaction Functions: ECB versus Bundesbank

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## *Abstract*

We estimate monetary policy reaction functions for the Bundesbank (1979:4-1998:12) and the European Central Bank (1999:1-2003:7). The Bundesbank regime can be characterised, both before and after German reunification, by an inflation weight of 1.2 and an output weight of 0.4. The estimates for the ECB are 1.2, and 1, respectively. Thus, the ECB, while reacting similarly to expected inflation, puts significantly more weight on stabilising the business cycle than the Bundesbank did.

**JEL:** E5

**Keywords:** Taylor rule, monetary policy, ECB, Bundesbank

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## 1. Introduction

In recent years, virtually all central banks in the industrialised countries have conducted monetary policy through market-orientated instruments designed to influence short-term interest rates (Borio 1997). Interest rate setting is commonly described in terms of the reaction to deviations of inflation and output from their targets (Taylor 1993). Gerlach and Schnabel (2000) estimate a Taylor rule for the euro area for the period 1990-1998 using synthetic euro area data<sup>1</sup>. They find that monetary policy can be well explained in terms of such a Taylor rule. Now that almost five years have passed since the start of EMU, enough observations have become available to perform a first assessment of the ECB's conduct of monetary policy using an empirically estimated Taylor rule.

The ECB was explicitly designed after the German Bundesbank in the Treaty of Maastricht. One may conjecture that if two institutions are based upon similar design features, in particular central bank independence and a policy focus on price stability, their interest rate setting rules may look alike, although the challenges facing the ECB may be different to those facing the Bundesbank in the past. In addition, formulating monetary policy as a compromise between the representatives of various countries could also result in a different interest rate behaviour. Thus, it will be interesting to compare estimated Taylor rules for the Bundesbank (see Clarida et al. 1998 and Bundesbank 1999) and the ECB.

Based on these considerations, the present paper estimates Taylor rules for both the ECB and the Bundesbank. For Germany, we use monthly data from the formation of the EMS onwards (1979:4 – 1998:12), explicitly taking into account the period after German unification (1990:8 – 1998:12). For the ECB, data from 1999:1 to 2003:7 is employed. It is then tested whether the ECB is behaving similarly to the Bundesbank with respect to interest rate setting.

## 2. Econometric Methodology

In the estimations of the Taylor rules, we follow Clarida et al. (1998).<sup>2</sup> Short-term nominal interest rates are modelled as a function of deviations of output from its trend and of inflation from its (constant) target. We assume that the central banks are forward looking and react to the expected one-year ahead inflation rate and the current output gap.<sup>3</sup> Finally, we allow for

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<sup>1</sup> Other studies estimating Taylor rules for the euro area based on synthetic euro area data are Mihov (2001), Doménech et al. (2002), Clausen and Hayo (2002), and Gerdesmeier and Roffia (2003).

<sup>2</sup> The utilised data are: Interest rate: Day-to-day rate, Output: Industrial production, Inflation: CPI growth rate. Data sources are IFS for Germany and the ECB web site. The output gap has been constructed using a Hodrick-Prescott filter. The typical endpoint problem of the filter is not an issue here, as both in the Bundesbank and in the EMU estimation we have at least 10 months of output gap data that do not enter the estimation process.

<sup>3</sup> Note that the output gap of the current period is not exactly known at the time of monetary policy decision making.

interest rate smoothing by including a lagged interest rate term in the Taylor rule specification<sup>4</sup>. Given these considerations, we estimate the following Taylor rule:

$$(1) \quad r_t = \rho r_{t-1} + (1-\rho) \alpha + (1-\rho) \beta \pi_{t+12} + (1-\rho) \gamma y_t + \varepsilon_t,$$

with:  $r$  = nominal short-term interest rate,  $\pi$  = inflation rate,  $y$  = output gap,  $\rho$  = degree of interest rates smoothing,  $\alpha$  = long-term equilibrium of nominal interest rate,  $\beta$  = inflation weight,  $\gamma$  = output weight,  $\varepsilon$  = error term.

A major problem when working with forward-looking and current variables is that they may be correlated with the error term, leading to biased estimates of the coefficients of interest. Therefore, these variables must be instrumented. In addition, the error term may experience non-normality, autocorrelation and heteroscedasticity, causing problems with respect to statistical inference. It is now common to use the General Method of Moments (GMM) estimator, as it accounts for endogeneity biases as well as non-spherical errors. The GMM estimator possesses excellent asymptotic properties, but may perform poorly in small samples (see the special issue of the *Journal of Economics and Business Statistics* 1996). Traditional instrumental variable estimators, which are special cases of the GMM estimator (Hayashi 2000), are more efficient, provided the necessary restrictions are fulfilled. In small samples, efficiency is of particular importance, and it therefore appears to be advisable to rely on a traditional instrumental variable estimator whenever possible.

Another general estimation problem is the choice of instruments. In time-series econometrics, it is easy to find instruments that fulfil the orthogonality conditions between regressors and error term. Typically, this assumption is tested using a test of the validity of over-identifying restrictions when there are more instruments than estimated coefficients (see Davidson and McKinnon 1993).

The use of weak instruments, i.e. instruments that do not contribute much to explaining the instrumented variable, can lead to substantial biases in both estimators and test statistics even in large samples (see, e.g., Hahn and Hausman 2003, Stock et al. 2002). Stock and Yogo (2003) propose a test of weak instruments based on the F-test value of the first stage regression in a two-stage least squares procedure. This test does not, however, solve the question of how to choose *specific* instruments, for example, which lags of a variable.

We address the instrument selection problem by adopting a recently developed automatic model selection algorithm called GETS (see Hendry and Krolzig 1999). GETS starts from a

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<sup>4</sup>Recent evidence shows that the explicit modelling of a lagged interest rate term is preferable to an autoregressive errors specification (Castelnuovo 2003).

general model and removes redundant instruments. While doing so, it searches all possible paths of the testing-down process and reports the most parsimonious model that does not violate a reduction test. Thus, the strongest instruments will be selected from a given choice of variables and their lags. This does not remove all arbitrariness, as, for instance, the researcher still needs to choose the potential instrumental variables and their maximum lag length, but it appears to be superior to the ad hoc methods typically employed in empirical research.

### 3. Empirical Estimation

Starting with the *Bundesbank*, instruments are being selected using the GETS reduction algorithm at a nominal 5% significance level based on a general model containing six lags of the potential instruments.<sup>5</sup> The Stock and Yogo (2003) test shows no indication of problems with regard to biases in the estimators. It is possible to reject the null hypothesis that the nominally sized test statistics at 5% significance level exceed an actual level of 15%. In contrast, there is evidence that the Clarida et al (1998) specification suffers from problems with weak instruments.<sup>6</sup>

Employing these instruments in traditional instrumental variable estimation leads to residuals that exhibit severe problems of non-normality, autocorrelation and heteroscedasticity. Since at least the non-normality problems are not easily corrected, we choose GMM as the appropriate estimation technique. An important event in recent German history is unification, with German Monetary Union (GMU) taking place on 1 July 1990. We split our sample at that date to see whether it has a noticeable impact on the German reaction function (see Table 1).<sup>7</sup> A comparison of the two columns referring to the Bundesbank shows that the estimates are similar. The lagged interest rate has exactly the same coefficient and the inflation coefficients are very close. Some differences occur for the output gap coefficient, which are not statistically significant when taking into account estimation uncertainty surrounding both coefficients ( $F(1, 231) = 1.89$ ).

Comparing these value to the ones Clarida et al. (1998) obtained for the period 1979:4 to 1993:12 shows only slight differences ( $\beta = 1.31$ ,  $\gamma = 0.25$ ). However, estimating the Clarida et

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<sup>5</sup> The resulting instrument set is: interest rate (lags: 1, 2, 3), inflation (lag: 6), growth rate of the effective real exchange rate (lags: 1, 4), output gap (lags: 1, 2, 3, 6), the growth rate of the oil price index in DM (lags: 1, 6), and the monthly growth rate of the money aggregate M3 (lag: 2).

<sup>6</sup> In their basic specification, Clarida et al. (1998) use 48 instruments (p. 1045, Table 1). In the first-stage regression for the inflation rate the Stock and Yogo-test can barely reject the hypothesis of a bias of 10% of the OLS bias and cannot reject the Null that the nominally sized test statistics of a 5% level does not exceed a level of 15%.

<sup>7</sup> One could estimate a Taylor rule from 1979:4 to 2003:7 using German data and then test for a break in 1999:1. This would not take into account, however, that the ECB uses aggregate European variables in its rule.

al. baseline model for the post-GMU period gives a coefficient of 0.74 for the inflation weight and 0.48 for the output gap. The drop of the coefficient on inflation below unity may be one indication of the weak instrument problem in their specification. As round summary values for the Bundesbank covering both periods, we propose a weight on inflation of 1.2 and an output weight of 0.4. These restrictions cannot be rejected using either pre- or post German Monetary Union data.<sup>8</sup>

Table 1: Estimates of reaction functions for the Bundesbank and the ECB

Variables	Bundesbank 79:4-90:6	Bundesbank 90:8-98:12	ECB 99:1-02:7
Interest rate <sub>t-1</sub> ( $\rho$ )	0.92** (0.018)	0.92** (0.015)	0.86** (0.033)
Inflation <sub>t+12</sub> ( $\beta$ )	1.21** (0.244)	1.25** (0.162)	1.12* (0.505)
Output gap <sub>t</sub> ( $\gamma$ )	0.43** (0.137)	0.32** (0.045)	1.03** (0.194)
Constant ( $\alpha$ )	3.64** (0.670)	2.56** (0.353)	1.38 (1.132)
No. of Observations	135	101	43
$\sigma$	0.372	0.154	0.157
$R^2$	0.976	0.996	0.966
Over-identifying restrictions test	$\text{Chi}^2(13) = 8.53$	$\text{Chi}^2(13) = 8.25$	$\text{Chi}^2(25) = 31.2$

*Notes:* \* (\*\*) indicates significance at a 5% (1%) level. Bundesbank estimates based on GMM. ECB estimates based on a traditional instrumental variable method. Standard errors for coefficient estimates are computed using the procedure by Newey and West (1987). The  $R^2$  is based on the short-run dynamic model. Diagnostic tests for the instrumental variable estimation of the ECB reaction function: Jarque-Bera normality test:  $\text{Chi}^2(2) = 0.65$ , LM autocorrelation test:  $\text{Chi}^2(2) = 2.35$ , ARCH test:  $\text{Chi}^2(1) = 0.68$ , White-heteroscedasticity test with cross-products:  $\text{Chi}^2(9) = 20.3^{**}$ , RESET(1) test:  $F(1, 37) = 1.11$ .

For the *ECB reaction function*, the data are from January 1999 to July 2003.<sup>9</sup> Again we select instruments based on the GETS algorithm.<sup>10</sup> Applying the Stock and Yogo (2003) test, we can reject the hypotheses that the biases in the estimators are larger than 10% and that the nominally sized test statistics at 5% exceed an actual level of 25%. This time we find there are

<sup>8</sup> Test results for the pre-GMU period: inflation coefficient against 1.2:  $F(1, 131) = 0.002$ , output coefficient against 0.4:  $F(1, 131) = 0.05$ , and joint:  $F(2, 131) = 0.03$ . Post-GMU period: inflation coefficient against 1.2:  $F(1, 97) = 0.09$ , output coefficient against 0.4:  $F(1, 97) = 2.88$ , and joint:  $F(2, 97) = 2.42$ .

<sup>9</sup> Given that we employ a one-year ahead inflation rate, the actual estimation period ends in July 2002.

<sup>10</sup> The resulting instrument set is: interest rate (lags: 1,2,4,6), inflation (lag: 1,3,5), growth rate of the effective real exchange rate (lags: 1,2,3,4,6), output gap (lags: 1,2,3,4,5,6), the growth rate of the oil price index in EUR (lags: 1,2,3,4,5), and the monthly growth rate of the money aggregate M3MA (lag: 2,3,4).

few problems with employing traditional instrumental variable techniques.<sup>11</sup> In the notes below Table 1, a battery of diagnostic tests are listed that show no problems except for evidence of heteroscedasticity. To avoid invalid inference, we employ, as in our estimates for Germany, robust standard errors.<sup>12</sup>

Regarding the actual estimates, we first note that the inflation coefficient is above unity. Second, it is quite close to the estimates we obtained for the Bundesbank. Statistically testing the ECB coefficient on inflation against 1.2, we cannot reject the hypothesis of equality ( $F(1,39) = 0.03$ ). On the other hand, the output weight estimated for the ECB is more than twice as large as the one found for the Bundesbank. Testing the ECB coefficient against 0.4 leads to a rejection of the hypothesis of equal size ( $F(1,39) = 10.5^{**}$ ).<sup>13</sup> This test does not take into account that there is uncertainty in the estimation of the German coefficient. Taking this uncertainty into account still leads to the conclusion that the ECB output coefficient is significantly larger than the one for the Bundesbank.<sup>14</sup>

## Conclusion

We estimate reaction functions for both the Bundesbank (1979:4 – 1998:12) and the European Central Bank (1999:1 – 2003:7) following the specification suggested by Clarida et al. (1998). We find that the Bundesbank reaction function can be characterised by an inflation weight of 1.2 and an output gap weight of 0.4 before and after German reunification. For the ECB we obtain a similar estimate of the inflation rate, and a unit coefficient on the output response, which is significantly higher. Thus, while the ECB reacts similar to the Bundesbank when it comes to inflation deviations, it shows a much stronger response to output deviations. Given the similar institutional design, a stronger response of the ECB to the output gap may reflect different preferences of the ECB compared to the Bundesbank or the new economic and political environment the ECB operates in.<sup>15</sup> Of course, whether these differences are temporary or permanent remains to be seen.

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<sup>11</sup> There is an outlier in September 2001, the time of the terrorist attack on New York, which we include in the list of instruments. Including the dummy in the reaction function itself leaves the other estimators almost unchanged.

<sup>12</sup> Normal standard errors are:  $\rho$  (0.038),  $\beta$  (0.730),  $\gamma$  (0.232), and  $\alpha$  (1.62).

<sup>13</sup> Eliminating the constant from the ECB equation raises the coefficient on inflation to 1.7 and makes it significant at a 1% level, while the output gap coefficient remains basically unchanged. However, from the point of view of both economic theory and history, it is implausible to have a zero long-run nominal interest rate.

<sup>14</sup> Testing the ECB output coefficient against the pre-GMU coefficient leads to a t-test value of 2.53. This is significant at the 5% level when performing a two-sided test and significant at a 1% level for a one-sided test. Against the post-GMU coefficient the value of the t-test is 3.57, which is significant at a 1% level.

<sup>15</sup> The coefficients in the Taylor rule depend upon policy preferences, economic structure, and shocks. In general it is impossible to identify the policy preferences from the estimated parameters of the reaction function. Cecchetti and Ehrmann (1999) discuss how to recover policy preferences from Taylor rule estimates under very restrictive assumptions.

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