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Kiel Working Paper No. 1221

Did the Revision of the ECB Monetary Policy Strategy Affect the Reaction Function?

by

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July 2004

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Abstract

In May 2003 the European Central Bank (ECB) announced the revision of its monetary policy strategy. Although the ECB stressed that the revision would not imply any fundamental change in their decisions, this remains to be verified empirically. Therefore, this paper tries to answer the question whether the strategy revision has induced a structural break in the ECB policy reaction function. To this end, we estimate several Taylor-type reaction functions and conduct a structural change analysis using both recursive parameter estimates and structural change tests. We find that the ECB has been following a stabilizing rule and that there is no clear-cut evidence in favor of a break after the revision, even though some signals of instability show up, particularly in June 2003.

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* We wish to thank Carsten-Patrick Meier for helpful comments.

** Financial support by the Borsa di Studio di Perfezionamento all´estero from Università Cattolica del Sacro Cuore is gratefully acknowledged.

1. Introduction

The aim of this paper is to assess whether the revision of the monetary policy strategy of the ECB has introduced any significant modification in observable monetary policy decisions. To this end, we specify and estimate a number of Taylor-type policy reaction functions. Based on this, we apply several stability tests to assess the structural stability of the estimated reaction functions over the course of the recent months.

In December 2002, after four years of conducting monetary policy for a new economic entity, the Governing Council of the ECB undertook an evaluation of its strategy and on 8 May 2003 announced the outcome of the strategy review, focusing mainly on the definition of price stability and on the structure of the two-pillar strategy. In particular, two remarkable changes over the initial strategy could be observed (ECB, 2003b). First, the ECB clarified that it intends to maintain inflation rates close to 2% p.a. This implies that it takes deflationary risks seriously by preventing the inflation rate from moving too close to zero. Second, money is not any longer explicitly assigned a "prominent role" in the conduct of monetary policy but is rather used to cross-check the results from an economic analysis of inflationary risks. By most commentators, this was viewed a downgrading of the monetary pillar (de Grauwe, 2003, Walton and Daly, 2003, Belke et al., 2003) even though Issing (2003) rejected this interpretation.

It is unclear, however, whether the May 2003 strategy revision in fact marks a change of the observable monetary policy actions. Obviously, a change of the inflation objective should be apparent in actual monetary policy. Gali et al. (2004) even speculate that the revision could be a first step towards a higher target range between 1% and 3% as adopted by other central banks. On the other hand, the ECB seemed to intend that the "close to 2%" announcement was rather a clarification regarding an acceptable lower bound of inflation rates than a change of its objective. Moreover, actual inflation rates had been too high between 1999 and 2003 to touch any lower bound, be it 0% or 1%. Therefore, it is questionable whether any changes in the conduct of monetary policy can actually be observed.

Similarly, the downgrading of the monetary pillar may be difficult to verify empirically even though, in principle, a step from monetary targeting towards inflation targeting may lead to a shift in policy actions. However, right from the start the ECB (1999a) rejected the notion that it followed a strict money-growth targeting. Empirically, the role played by M3 in the practice of the ECB is rather dubious (Begg et al., 2002, Gali, 2002, Favero et al., 2000, Svensson, 2000, von Hagen, 1999). Therefore, it is again unclear whether the announced revision has lead to any empirically detectable policy shift.

The preceding discussion suggests that it remains to be analysed empirically whether the policy actions of the ECB have changed due to the revision. Moreover, there are now quite a few observations after the suspected break date available so that the power of statistical test applied to this question should be non-negligible. We analyze the stability issue with the help of estimated ECB policy reaction functions and a structural change analysis. A number of authors[1] has estimated Taylor-type reaction functions for the Euro area but they rarely assess the structural stability of their estimates. Moreover, pre-EMU and EMU data are typically merged to obtain large samples but the implicit assumption of structural stability during the transition from 11 independent central banks to the ECB is hard to defend and may be particularly problematic when analyzing stability in another part of the sample. Therefore, we estimate the ECB reaction function with EMU data since January 1999 only.

The structural change analysis is performed in two steps. First, we compute recursive estimates to provide a rough signal of possible parameter instability. While this approach may yield suggestive results, it does not lend to interpretation in terms of statistical significance. Therefore, we additionally calculate stability tests. Their choice is guided by two considerations. First, they should be applicable to relatively short samples, and second, they should be designed to test for instability at the sample end. Hence, we choose a prediction test proposed by Dufour et al. (1994) and an end-of-sample instability test introduced by Andrews (2003).

The paper is organized as follows. In the next section the policy reaction function is outlined. Section 3 describes the data and the estimation methodology, while the estimation results are presented in Section 4. A structural change analysis is provided in Section 5. Finally, Section 6 concludes.

2. The policy reaction function specifications

Since Taylor (1993) it is typically assumed that the behaviour of the central bank can be characterized by the following policy reaction function:

[1] These are, *inter alia*, Gerlach and Schnabel (2000), Gerdesmeier and Roffia (2003), Surico (2003), Carstensen (2003), Gerlach-Kristen (2003) and Gali et al. (2004). For an overview of the empirical literature see, e.g., Sauer and Sturm (2003).

$$i_t = r^* + \pi_t + 0.5(\pi_t - \pi^*) + 0.5\widetilde{y}_t = r^* - 0.5\pi^* + 1.5\pi_t + 0.5\widetilde{y}_t, \quad (1)$$

where i_t is the policy interest rate, r^* the equilibrium real interest rate, y_t output, y_t^* potential output, $\widetilde{y}_t = y_t - y_t^*$ the output gap, π_t the rate of inflation, and π^* the inflation target. In this formulation, the interest rate is the policy instrument and depends on both inflation and the output gap. As shown by Svensson (1999b), this does not preclude that the ECB follows a strict inflation targeting if there is a structural relationship like a Phillips curve between inflation and the output gap. In this case, the output gap helps predicting inflation and therefore shows up in the reaction function which is in line with the formulation of the second pillar of the monetary policy strategy of the ECB (1999a).

Instead of taking them as given, many authors have tried to estimate the weights attached to inflation and the output gap. The general idea is to estimate the following equation:

$$i_t = r^* + \pi_t + \beta_1\widetilde{y}_t + (\beta_2 - 1)(\pi_t - \pi^*) + \varepsilon_t = \beta_0 + \beta_1\widetilde{y}_t + \beta_2\pi_t + \varepsilon_t, \quad (2)$$

where β_0 is the constant term $r^* - (\beta_2 - 1)\pi^*$, β_1 is the coefficient that captures the weight of the output gap, β_2 is the coefficient that captures the weight of inflation, and ε_t is an error term that captures any *ex post* deviations from the reaction function. Strictly speaking, the reaction function (2) is an implicit instrument rule in the terminology of Svensson (1999a), because contemporaneous variables cannot be observed at the time of the interest rate decision.[2]

With regard to the expected size of the coefficients, an important empirical question relates to the weight of inflation. Given that the private decisions of consumption and investment are affected by the real interest rate, β_2 must be large enough to raise the real interest rate when a rise in inflation occurs which is the so-called Taylor principle. From (2) it is obvious that a sufficient condition for this to hold is $\beta_2 - 1 > 0$ and, hence, $\beta_2 > 1$.[3] Otherwise, the rational-expectations equilibrium may be undetermined and the economy may drive into a self-fulfilling inflation spiral (Woodford, 2001). Regarding the weight of the

[2] This is not only due to publication lags but also because an interest rate decision for period *t* is made before this period actually starts. Many authors have estimated explicitly forward-looking reaction functions. However, our small sample at hand prevents this because using, e.g., a one-year ahead inflation rate would not only reduce the sample by one year but also preclude the structural change analysis at the sample end.

[3] Note that if $\beta_1 > 0$ the necessary condition may be less strict because a rise in inflation goes hand in hand with a rise of the output gap which leads to an additional interest rate response, see Woodford (2001).

output gap, we expect $\beta_1 \geq 0$ so that monetary policy has a stabilizing (or no) effect on output.

Since central banks seem to adjust interest rates gradually, probably in order to avoid disturbing markets, we follow Clarida et al. (1998) who assume a partial adjustment of the actual rate, i_t, towards its target level, i_t^*. In our context, this interest rate smoothing can be modeled as

$$i_t = \rho i_{t-1} + (1-\rho) i_t^*, \tag{3}$$

where ρ is the smoothing parameter. This results in the following compact form of the reaction function to be estimated:

$$i_t = \rho i_{t-1} + (1-\rho)(\beta_0 + \beta_1 \tilde{y}_t + \beta_2 \pi_t + \beta_3 x_t) + \varepsilon_t, \tag{4}$$

where x_t is an additional exogenous variable with weight β_3 that enters the ECB reaction function. In this study, three different reaction functions are estimated. In addition to the baseline model which resembles a standard Taylor rule because it only includes inflation and the output gap, we consider either money growth or the nominal effective exchange rate as additional arguments for the ECB reaction function. We include money growth to model the first pillar of the monetary policy strategy of the ECB which emphasizes the prominent role of M3 growth for interest rate decisions. This may reflect the leading indicator properties of money growth for inflation (Altimari, 2001) and the output gap (Coenen et al., 2001). Alternatively, we include the exchange rate to model foreign influences on the ECB interest rate decisions. According to the second pillar of its monetary policy, the ECB claims to pay attention to a broad set of economic variables that may help to assess the presence of threats to price stability. While it is not clear whether central banks directly react and should react to exchange rate changes (Taylor, 2001), the ECB might have been particularly tempted to counteract devaluations in the first years of EMU in order to establish the notion of a strong Euro as an equivalent successor of the German Mark. From a theoretical perspective, a direct influence of exchange rate changes in the instrument rule can pay off in terms of reduced inflation variance (Ball, 1999, Taylor, 1999).

3. The data and the estimation procedure

We use two different approaches to estimate the ECB reaction function. Following most of the literature, we use ex-post realized data and apply the generalized method of moments

(GMM). To cross-check the results, we also use survey data which are known at the time of an interest rate decision and allow us to apply ordinary least squares (OLS). In this section, we describe the data and the estimation procedures.

3.1. The GMM approach

The GMM approach is essentially an instrumental variables estimation of Equation (4) and is necessary because at the time of an interest rate decision, the ECB cannot observe the *ex post* realized contemporaneous right-hand side variables in (4). Therefore, it bases its decisions on an information set which comprises lagged variables only. The weighting matrix in the objective function is chosen in order to allow the GMM estimates to be robust to possible heteroskedasticity and serial correlation of unknown form in the error terms.[4]

With regard to the choice of the instruments, they need to be predetermined at the time of an interest rate decision, i.e., dated *t*-1 or earlier, and they should help predict the at time *t* unobserved contemporaneous variables, in particular inflation. Therefore, we include the first four lags of the nominal interest rate, inflation, the output gap, money growth, and the real effective exchange rate. The former three variables are typically used as instruments in related work (Sauer and Sturm, 2003, Gerdesmeier and Roffia, 2003). We add money growth because, due to the first pillar, it obviously plays a role in forming inflation expectations by the ECB, and the nominal effective exchange rate for the reasons laid out above. The choice of a relatively small number of lags for the instruments is intended to minimize the potential small sample bias that may arise when too many overidentifying restrictions are imposed.[5] To confirm that we have chosen an appropriate instrument set, we run a first stage regression of inflation on the instrumental variables and perform an *F*-test for their joint significance.

A second important property of the instrumental variables is their exogeneity with respect to the central bank decisions and, hence, their uncorrelatedness with the disturbances ε_t which reflect deviations from the policy rule that are unpredictable *ex ante*. To test this, we perform a standard *J*-test for the validity of the overidentifying restrictions (Hansen, 1982).

[4] We apply a four lags Bartlett window to account for possible autocorrelation.

[5] We also used different sets of instrumental variables and a larger number of lags. However, the magnitude of the coefficients and the general conclusions remained unaffected.

3.2. *The ex post data set*

The GMM estimation is conducted for the Euro period from January 1999 to February 2004 on aggregated Euro area data. To generate lagged instruments and estimate the output gap, earlier synthetic Euro area data are used. All data are monthly and seasonally adjusted and have been collected from Eurostat and from the ECB database. The variables are defined as follows: inflation is measured as the annualized rate of change of the Harmonized Index of Consumer Prices (HICP), the Euro Overnight Index Average (EONIA) and the 3-month EURIBOR are used as the short-term policy variable, the monetary aggregate M3 is constructed using the data on seasonally adjusted month-end stocks and flows from which annual rates of change are calculated. As exchange rate variable we used the annual growth rate of the euro nominal effective exchange rate calculated for a broad group of currencies.

With respect to the choice of the output gap measure, we use a "pseudo" real-time approach since true real-time data are not available for the Euro area. The pseudo real-time output gap in period t, $\tilde{y}_t$, is calculated from the *ex post* revised log industrial production index excluding constructions using only observations up to period t. By this we mimic the non-availability of information beyond t which characterizes the true real-time situation. Our approach is motivated by a finding of Clausen and Meier (2003) that using true and pseudo real-time output gaps for Germany yield very similar estimated Bundesbank reaction functions while typical *ex post* output gaps calculated from all data available to the researcher lead to quite different results. Technically, we employ a recursive Hodrick-Prescott Filter with smoothing parameter set at $\lambda = 129600$ as advocated by Ravn and Uhlig (2002). In order to obtain reliable trend estimates, we use observations from January 1993 onwards.

3.3. *The survey data*

In addition to the GMM approach, we estimate the reaction function using survey data. Following Sauer and Sturm (2003), we use the deviation of the composite EU Economic Sentiment Indicator (ESIN)[6] from its average over the relevant time period. Since the ESIN is an index number, it lacks a natural scale comparable with the output gap. Therefore, we centred and rescaled the ESIN gap so that is has the same mean and the same variance as the

[6] The EU ESIN is published by the European Commission on monthly basis and is a weighted average of an industrial confidence indicator, a services confidence indicator, a consumer confidence indicator, a retail trade confidence indicator and a construction confidence indicator. It is available one or two months before the industrial production index.

output gap. Hence, the patterns of the two indicators are rather similar, although the short-run volatility of the ESIN is smaller than that of the output gap and the former seems to lead the latter.

Concerning the inflation measure, instead of relying on the statistical releases of the HICP, we also use survey results to get an idea of expected inflation developments. Every month, Consensus Forecasts publishes inflation forecasts based on a poll of a group of international forecasters. As these figures are not, say, 12-month-ahead inflation forecasts but expected inflation rates for single years, a weighted average of the forecast for the current and the following year is taken as a proxy for the 12-month-ahead forecast of inflation[7]. This inflation forecast measure is less volatile and seems to be a leading indicator of actual inflation. It should be noted that a reaction function based on this indicator has a more forward-looking element than a reaction function based on contemporaneous ex post inflation. All data are plotted in Figure 1.

As the indicators are available at the time of the interest rate decisions, OLS is sufficient to obtain consistent estimation results. This has the advantage that the results are unaffected by the choice of tuning parameters (instrument lag, nonparametric covariance estimator) which are used for GMM estimation. Nevertheless, we also apply GMM and compare the results, taking OLS as a simple robustness check of the GMM results.

4. The estimation results

We first present the estimation results of the baseline model where the interest rate setting depends on the lagged interest rate, the output gap and inflation. Subsequently, we explore two augmented specifications where the money growth rate and the exchange rate change are taken as additional regressors.

4.1. The baseline policy reaction function

As our baseline model, we consider a standard policy reaction function with inflation and the output gap as explanatory variables. Consequently, the reaction function (4) can be re-written as:

$$i_t = \rho i_{t-1} + (1-\rho)\beta_0 + (1-\rho)\beta_1 \tilde{y}_t + (1-\rho)\beta_2 \pi_t + \varepsilon_t . \qquad (5)$$

[7] The weights are $x/12$ for the x remaining months in the current year and $(12-x)/12$ for the following year's forecast. See also Smant (2002).

It is estimated in reduced form

$$i_t = \rho i_{t-1} + \delta_0 + \delta_1 \tilde{y}_t + \delta_2 \pi_t + \varepsilon_t \tag{6}$$

from which the structural parameters, β_0, β_1 and β_2 are calculated back, using the relations $\delta_0 = (1-\rho)\beta_0$, $\delta_1 = (1-\rho)\beta_1$ and $\delta_2 = (1-\rho)\beta_2$. The standard errors of the structural parameters are obtained by means of the delta method.

The baseline model is estimated in four different specifications. For specifications S1 and S2 we take the overnight rate and the 3-month Euribor, respectively, as policy variables and the pseudo real-time output gap and actual inflation as explanatory variables. For specifications S3 and S4 we again take the overnight rate and the 3-month Euribor, respectively, as policy variables, but the survey indicators as explanatory variables. In addition to GMM, we apply OLS as a means of cross-checking. The estimation results are presented in Table 1.

When using the overnight rate as policy instrument and applying GMM (column 3), an estimated smoothing parameter $\hat{\rho} = 0.96$ indicates a considerable degree of persistence in the interest rate dynamics. The coefficients of the output gap and inflation in the reduced form equation (6) represent the partial effects conditional on a given lagged interest rate and are estimated highly significant as 0.07 and 0.08, respectively. This means that the ECB immediately responds to a one percentage point rise of the output gap by raising the interest rate by 7 basis points and to a one percentage point rise of inflation by raising the interest rate by 8 basis points. The validity of the instruments is confirmed by a highly significant first-stage regression *F*-statistic and an insignificant *J*-statistic.

The long-run weight of inflation turns out to be $\hat{\beta}_2 = 1.89$. As this is larger than one, the policy response to an increase in inflation is in line with the Taylor principle. Thus, the ECB appears to follow a stabilizing course, in the sense that nominal policy rate changes are large enough to affect real short term interest rates in the same direction.

The long run weight of the output gap is calculated as $\hat{\beta}_1 = 1.70$. This supports the evidence shown by Faust et al. (2001) who argue that the ECB has placed a relatively high weight on the output gap compared to the Bundesbank. Different results are obtained by Fourçans and Vranceanu (2002), who find the ECB to react strongly to variations in the inflation rate and much less to output variations, and by Gerdesmeier and Roffia (2003) and Ullrich (2003), whose estimated output gap weight is below unity.

The conclusions remain largely unaffected when we estimate specification S2, i.e., we replace the overnight rate with the 3-month Euribor (column 5). However, especially the weight of inflation deteriorates to 1.10 which is only slightly above unity. Still, the large weight of the output gap should ensure that the reaction of the ECB to a rise of inflation is strong enough to remain safely within the stability region and, hence, satisfy the Taylor principle. Compared to specification S1, the *J*-statistic deteriorates a little. While it is still insignificant, this may indicate that the overnight rate is better suited as the policy variable than the 3-month Euribor.

When applying OLS instead of GMM to specifications S1 and S2 (columns 4 and 6), only minor changes can be observed, particularly when estimation uncertainty is taken into account. In addition, we investigated the sensitivity of the GMM results to different sets of instrumental variables. The inclusion of the long-short interest rate spread and of the commodity price inflation as well as the exclusion of the money growth rate do not lead to important changes in the magnitude of the estimated coefficients.

When using the survey indicators as proxies for the expected output gap and expected inflation, the magnitude of the coefficients changes appreciably. Taking the overnight rate as policy instrument (specification S3) and applying GMM, we obtain a lower smoothing parameter, $\hat{\rho} = 0.88$, a lower output weight $\hat{\beta}_1 = 0.71$ and a higher inflation weight $\hat{\beta}_2 = 2.94$ than in specification S1 where the pseudo real-time output gap and inflation are used as explanatory variables. Similar changes occur for specification S4 compared to S2, and for OLS estimation. From all results, the ECB appears to follow an even more stabilizing policy rule concerning inflation, but places a smaller weight on the output gap than found in the previous specifications. However, the higher weight of inflation can at least partially be explained by the fact that the Consensus Forecast for inflation has a smaller variance than actual inflation, see Figure 1, which is a typical property of rational forecasts. Everything else unchanged, replacing actual with forecasted inflation should then yield a larger parameter estimate. Apart from this technical aspect, the results indicate that the ECB places a much larger weight on inflation than on output, once we use variables which are known at the time of an interest rate decision. Taking the estimation uncertainty into account, the OLS results for S3 and S4 are not even far away from the Taylor weights of $\beta_1 = 0.5$ and $\beta_2 = 1.5$. This implies that it is difficult to argue that the ECB is less concerned with inflation than other central banks as early commentators did (Faust et al., 2001, Gali, 2002). Instead, it rather confirms the results obtained by Gerdesmeier and Roffia (2003) and Begg et al. (2002) that the ECB behaves not too differently from the Fed.

4.2. Augmented policy reaction functions

So far, we have analyzed reaction functions which only include measures of inflation and the output gap as explanatory variables. In the following, we augment this baseline model with the annual M3 growth rate (model 2) and the nominal effective exchange rate (model 3). These variables might have played an important role in the interest rate setting of the ECB and can thus have some explanatory power in addition to the baseline variables. Both for model 2 and model 3, we again estimate the same four different specifications as for the baseline model. In specifications S1 and S2 we take the overnight rate and the 3-month Euribor, respectively, as policy variables and the pseudo real-time output gap and actual inflation as explanatory variables. In specifications S3 and S4 we again take the overnight rate and the 3-month Euribor, respectively, as policy variables, but the survey indicators for the output gap and inflation.

The estimation results are presented in Table 2. Adding money growth to the baseline variables yields model 2 which has a stronger degree of interest rate smoothing than before. In specification S2, we even obtain a point estimate of $\hat{\rho} = 1.01$ which implies that the model is explosive and the long-run weights are not meaningful anymore. Similarly, in specification S1, the estimate of 0.97 is so near to 1 that the estimation uncertainty of the long-run weights becomes quite large. From an economic point of view, this evidence can be interpreted as follows. Since ρ captures the influence of the lagged interest rate on the current interest rate decision, i_{t-1} becomes more and more important as ρ tends to one. Consequently, the relative importance of the other explanatory variables diminishes. In the extreme case $\rho = 1$ they are not suitable anymore to explain the long run patterns of the policy variable. Smoothing parameter estimates a bit more away from 1 are obtained in specifications S3 and S4 where the survey indicators replace the output gap and inflation. However, the weight of the money growth rate turns out insignificant in any case so money does not seem play an important own role in the interest rate setting of the ECB.

Alternatively, we add the log changes of the nominal effective exchange rate to the baseline variables (model 3). The results of the four specifications are again shown in Table 2. Even though the coefficient of the exchange rate is relatively small compared to the ones of the other explanatory variables, it is highly significant and has the expected sign. As discussed in Taylor (2001), an appreciation (a rising exchange rate) leads to a relaxation of monetary policy. Moreover, our point estimates are in the range analyzed by Taylor (1999).[8]

Regarding the remaining parameters, a comparison of model 3 with model 1 reveals that the weights of the output gap and inflation have decreased, in some cases dramatically. Let us first consider specification S1 where the overnight rate is taken as the policy instrument. The weights of the output gap and inflation are estimated 0.69 and 1.37 in model 3 opposed to 1.70 and 1.89, respectively, in model 1. We may thus conclude that adding the exchange rate leads to weights much more in line with the original Taylor rule. However, this result does not carry over to all other specifications. In specification S2, where the 3-month Euribor is taken as the policy instrument, we obtain an inflation weight of 0.67 which violates the Taylor principle. Specifications S3 and S4, which use the survey indicators instead of the output gap and inflation, look more sensible because inflation has a weight far above unity. Finally, the smoothing parameter is estimated considerably smaller than in model 2 and slightly so than in model 1.

5. Structural change analysis

Having estimated a number of reaction functions under the implicit assumption of structural stability, it is now analyzed whether the revision of the ECB monetary policy strategy announced in May 2003 has led to any perceivable parameter changes in the policy reaction function. In order to detect the presence of a structural break, a recursive analysis and several stability tests are performed. We start from the baseline sample 1999:1 to 2002:12 during which we assume a stable reaction function. This choice seems sensible for two reasons. First, it allows us to have 4 full years of observations available for the pre-break estimation. Second, January 2003 seems to be a suitable starting point for the search of a break point because the process of evaluation of the monetary policy was first announced in December 2002 to start in the following year (ECB, 2002) and, hence, a modification of the reaction function could have been implemented before the public announcement in May 2003.

[8] However, note that the reaction function proposed by Taylor (1999) is not strictly comparable to ours because he includes both the change and the level of the exchange rate with weights –0.15 and –0.1, respectively.

5.1. *Recursive Analysis*

Although it is not strictly interpretable as a test, a recursive analysis is a valuable tool because it provides a *signal* of possible instability in the parameters of the reaction function. To this end, equation (4) is first estimated on the baseline sample. Then, the parameters are re-estimated recursively on a sample including one more observation for each further step. The evolution of the estimates over the sample 2003:1 to 2004:2 is presented in the first three rows of Figures 2 to 5. The graphs display the estimated parameters $\hat{\rho}$, $\hat{\beta}_1$ and $\hat{\beta}_2$ together with 95% confidence intervals.

The left column of Figure 2 shows the estimation results for the baseline model in specification 1, i.e., with the overnight rate as policy variable, and with the pseudo real-time output gap and inflation as explanatory variables. The recursive analysis reveals that since the beginning of 2003, the smoothing parameter has started to increase from roughly 0.89 to 0.96. This induces a widening of the confidence intervals of the other weights because the stationarity border comes close. The weight of the output gap is increasing particularly strongly from about 0.7 in 2003:1 to 1.7 in 2004:2 while the weight of inflation increases somewhat less from 1.3 to 1.9. This could imply that the ECB has in fact changed its reaction function during this time, placing relatively more weight on output than before.

Similar patterns can be found from the recursive estimates of the baseline model in specification S2, where the 3-month Euribor replaces the overnight rate (Figure 2, right column). However, while the estimates for the smoothing parameter and the weight of the output gap increase, the inflation weight stays rather constant at 1.1.

An interesting result emerges from comparing GMM and OLS estimates for specification S3, displayed in Figure 3. As the pseudo real-time output gap and inflation are replaced by the survey indicators, OLS is a consistent alternative to GMM. The GMM estimates (left column) and the OLS estimates (right column) of the smoothing parameter and the weight of the output gap show the same upward trend found in specification S1, even though the absolute changes are smaller. However, the weight of inflation stays constant at 2.4 until 2003:9 and increases only afterwards to 2.9 when estimated by GMM while it remains at 2.4 over the full period when estimated by OLS. From this we may conclude that the estimation strategy has an influence on the stability results.

At the first sight, the recursive estimates for model 2, where money growth is added as an explanatory variable, seem to support the hypothesis of stability (Figure 4). However, this is mainly due to the very large confidence intervals which reflect the near non-stationarity of the interest rate process. Interestingly, the money growth coefficient is insignificant over the whole time span which indicates that money growth did not have an important influence on interest rate decisions even before the strategy revision was undertaken. If, alternatively, the exchange rate is added as an explanatory variable, we obtain quite stable estimates for the weights of the output gap and inflation (Figure 5). In contrast, the weight of the exchange rate is decreasing over time while the smoothing parameter shows a similar upward drift as found before.

In sum, the recursive analysis shows some possible signs of instability in the estimated coefficients of the ECB reaction function over the sample 2003:1 to 2004:2. There is no specification which looks unambiguously stable. Nevertheless, it is important to point out that this exercises provides only a *signal* of instability but does not enable us to assess whether the changes in the estimated parameters are statistically significant. Therefore, in a second step, we compute statistical tests to clarify this issue.

5.2. Structural Break Tests

There are a number of structural break tests for GMM estimation available in the literature, some of which take the break point as known and some of which treat it as endogenous. In our context, the key feature of the testing problem is that the number, m, of observations in the period after the potential change is relatively small. This prevents the use of tests which require both the pre-break and the post-break sample to be large, or which can only be applied to a trimmed sample since we are explicitly interested in the sample end.

Following Dufour et al. (1994), we perform a predictive test for structural stability suitable when the model is structurally stable during a baseline sample but possibly instable thereafter. This test is a generalization of the predictive Chow test to the GMM environment. Its most important feature is that it does not require the separate estimation of a post-break equation. Given the small number of observations after May 2003, this is particularly important in our study. The first step of the test procedure is to estimate the parameters of the ECB reaction function in the baseline sample which are then used to obtain a fitted interest rate series for the prediction sub-sample. The standardized prediction errors, $\tilde{\varepsilon}_t$, follow an iid standard normal distribution if a number of strong assumptions such as stationarity, normality, homoskedasticity and absence of autocorrelation hold. Prediction errors outside a, say, 95% interval indicate structural instability in specific periods. In addition, as an overall test statistic we calculate the sum of the squared prediction errors, U, which is, under the null hypothesis of stability, χ^2-distributed with degrees of freedom equal to the number of observations, m, in the prediction sub-sample.

The graphs in the fourth row Figures 2 to 5 show the standardized prediction errors while the U-statistics are reported in the lower panels of Tables 1 and 2. A first interesting observation is that the prediction errors are generally negative, which implies that the model overpredicts the actual interest rate development, but that this overprediction is mostly not significant. Only in June 2003, the month after the announcement of the policy revision, there is a significantly negative prediction error in all models and specifications. This is a strong indicator for some kind of instability in this period. On the other hand, it is well-known from the literature that using sequential tests like this when the true breakpoint is not exactly known may lead to size distortions and overrejection of the true null hypothesis of stability (see, e.g., Maddala and Kim, 1998). Therefore, we also report the U-statistic which tests the null of stability against the alternative that a break occurred after January 2003. While this test may have low power to detect instability in May or June 2003, it should at least be correctly sized. The results confirm the view that a single spike in the prediction errors does only lead to an overall rejection of stability if the other prediction errors are also rather large. For example, specifications S1 and particularly S2 of the baseline model are indeed unstable while specification S3 is not. The same observation can be made for the other models. Comparing all test results, the most important finding is that the U-test generally rejects stability when the pseudo real-time output gap and inflation are used as regressors but accepts stability when the survey indicators are used.

However, the preceding tests require strong distributional assumptions. If they are violated, inference may be misleading. To at least partially overcome this problem we implement the test of Andrews (2003) for end-of-sample structural instability. Its motivation is similar to the one in Dufour et al. (1994), but the critical values are obtained from parametric sub-sampling which gives asymptotically correct results under much weaker distributional assumptions. Parametric sub-sampling means that the distribution of the test statistic is derived from sequentially applying the test to the supposedly stable estimation sub-sample. This implies that the critical values can differ considerably from specification to specification. This approach is problematic if the estimation sub-sample is rather short as in our case because then the critical values cannot be estimated very precisely. As the time of the possible break is not exactly known (even though May or June 2003 are plausible candidates), we perform this test for each observation in the sample 2003:1 to 2004:2 and present the corresponding p-values in the graphs in the bottom rows of Figures 2 to 5. The test for instability after June 2003 is also reported in the bottom row in Tables 1 and 2. Now the hypothesis of stability cannot be rejected at any conventional significance level, even though some weak signals of instability corresponding to the June 2003 observation are also confirmed by the Andrews test.

We can thus summarize that the finding of instability heavily depends on the test under use. Assuming iid normal disturbances, we are able to strongly reject stability in a number of cases. However, relaxing this assumption in favour of mere stationarity, we only obtain weak–and insignificant–signals of instability. Therefore, for the moment we prefer those specifications which take the survey indicators as explanatory variables because they are found stable by both approaches. This confirms a finding by Gali et al. (2004) who recommend a similar specification.

6. Concluding remarks

In this paper we estimated a Taylor-type reaction function of the European Central Bank for the first five years of EMU using either the overnight rate or the 3-month Euribor as policy instruments. The results of our analysis indicate that the question whether the revision of the ECB monetary policy strategy in May 2003 has introduced a significant change in observable monetary policy decisions cannot be denied unambiguously for all specifications. In particular, the following conclusions can be drawn.

First, we found plausible specifications for the ECB policy reaction function, which indicates that the ECB has been following a stabilizing rule over its first years of existence. This holds no matter whether we used ex-post available data for the output gap and inflation or survey indicators which are available to the ECB at the time of an interest rate decisions. As a particularly interesting side-aspect, we were not able to find a significant impact of money growth on the interest rate decisions.

Second, from a recursive analysis there are some signs of instability, particularly in June 2003, directly after the announcement of the policy strategy revision. However, it is difficult to verify this by means of a statistical test. Using ex-post available data for the output gap and inflation as explanatory variables leads to conflicting test results indicating at least that stability cannot be assumed without any doubt. Using instead survey indicators as explanatory variables leads to much less evidence of instability and should, therefore, be preferred.

Overall, it may be fair to conclude that at the moment there is not enough evidence to reject the hypothesis of structural stability of the ECB policy reaction function, even though some minor doubts remain. This implies that the obvious downgrading of the monetary pillar in the policy strategy did not result in an abrupt change of actual policy. Probably, it was rather an attempt to reconcile words with deeds than anything else.

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Tables

Table 1: Estimates of the baseline policy reaction function

Explanatory variable	Coefficient	S1		S2		S3		S4	
		GMM	OLS	GMM	OLS	GMM	OLS	GMM	OLS
Interest rate (t–1)	ρ	0.957	0.947	0.946	0.936	0.884	0.895	0.867	0.885
		(0.015)	(0.026)	(0.017)	(0.026)	(0.015)	(0.030)	(0.028)	(0.032)
Intercept	δ_0	0.000	0.001	0.001	0.001	–0.001	0.000	0.001	0.001
		(0.001)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)
Output gap	δ_1	0.073	0.071	0.074	0.070	0.082	0.081	0.088	0.082
		(0.006)	(0.012)	(0.007)	(0.013)	(0.006)	(0.012)	(0.009)	(0.013)
Inflation	δ_2	0.081	0.085	0.060	0.074	0.342	0.248	0.259	0.199
		(0.027)	(0.046)	(0.026)	(0.044)	(0.053)	(0.096)	(0.076)	(0.095)
Intercept	β_0	0.008	0.012	0.021	0.019	–0.012	–0.002	0.005	0.009
		(0.011)	(0.014)	(0.006)	(0.013)	(0.006)	(0.011)	(0.004)	(0.011)
Output gap	β_1	1.696	1.340	1.359	1.094	0.709	0.767	0.659	0.714
		(0.552)	(0.605)	(0.391)	(0.567)	(0.061)	(0.186)	(0.089)	(0.175)
Inflation	β_2	1.889	1.611	1.099	1.155	2.949	2.354	1.947	1.724
		(0.608)	(0.764)	(0.256)	(0.738)	(0.343)	(0.630)	(0.211)	(0.612)
F-statistic		19.158		19.420		43.700		41.970	
(p-value)		(0.000)		(0.000)		(0.000)		(0.000)	
J-statistic		10.954		12.546		11.284		12.338	
(p-value)		(0.859)		(0.766)		(0.841)		(0.779)	
U-statistic		28.206	29.323	53.072	46.666	19.409	17.873	29.306	25.173
(p-value)		(0.013)	(0.009)	(0.000)	(0.000)	(0.150)	(0.213)	(0.009)	(0.033)
Andrews test		13.663	10.238	9.18	11.414	14.773	11.365	23.196	19.424
(p-value)		(0.689)	(0.333)	(0.689)	(0.778)	(0.156)	(0.222)	(0.556)	(0.667)

Notes: Specifications S1 and S3 take the overnight rate as policy instrument, specifications S2 and S4 use the 3-month Euribor. Specifications S1 and S2 use the pseudo real-time output gap and the inflation rate as explanatory variables, specifications S3 and S4 use the survey indicators. All estimated parameters are displayed with standard errors in brackets below. The GMM instrument set includes lags 1 to 4 of the interest rate, inflation, the output gap, the money growth rate and the nominal effective exchange rate. The F-statistic refers to the first stage regression and is F-distributed. The J-statistic of the Hansen test for overidentifying restrictions is χ^2-distributed. The U-statistic refers to the Dufour et al. (1994) predictive test for structural stability. Under the null hypothesis of stability, it is χ^2-distributed with degrees of freedom equal to the number of observation in the prediction sub-sample (14). The value for the Andrews test refers to June 2003 as the time of possible break. The p-value is obtained from parametric sub-sampling.

Table 2: Estimates of the augmented policy reaction function

Explanatory variable	Coefficient	Model 2: money growth				Model 3: exchange rate			
		S1	S2	S3	S4	S1	S2	S3	S4
Int. rate (–1)	ρ	0.970	1.011	0.925	0.956	0.938	0.893	0.880	0.861
		(0.019)	(0.029)	(0.022)	(0.028)	(0.012)	(0.009)	(0.014)	(0.018)
Intercept	δ_0	–0.001	–0.004	–0.004	–0.006	0.001	0.003	–0.001	0.001
		(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.000)	(0.001)	(0.001)
Output gap	δ_1	0.082	0.102	0.104	0.128	0.043	0.014	0.058	0.017
		(0.014)	(0.014)	(0.016)	(0.011)	(0.005)	(0.007)	(0.011)	(0.010)
Inflation	δ_2	0.070	0.008	0.267	0.164	0.085	0.072	0.322	0.215
		(0.028)	(0.035)	(0.060)	(0.068)	(0.026)	(0.021)	(0.051)	(0.058)
Money growth	δ_3	0.016	0.061	0.044	0.089				
		(0.020)	(0.024)	(0.024)	(0.020)				
Exchange rate	δ_4					–0.008	–0.018	–0.006	–0.018
						(0.002)	(0.001)	(0.003)	(0.002)
Intercept	β_0	–0.025	0.310	–0.054	–0.133	0.013	0.024	–0.009	0.010
		(0.062)	(0.644)	(0.031)	(0.105)	(0.008)	(0.004)	(0.007)	(0.005)
Output gap	β_1	2.770	–8.937	1.395	2.936	0.695	0.132	0.481	0.119
		(2.106)	(21.594)	(0.548)	(1.929)	(0.171)	(0.063)	(0.112)	(0.066)
Inflation	β_2	2.350	–0.708	3.575	3.766	1.368	0.670	2.691	1.542
		(1.223)	(4.684)	(0.508)	(1.272)	(0.407)	(0.185)	(0.391)	(0.295)
Money growth	β_3	0.525	–5.321	0.594	2.028				
		(0.933)	(11.547)	(0.467)	(1.649)				
Exchange rate	β_4					–0.128	–0.167	–0.053	–0.132
						(0.026)	(0.011)	(0.025)	(0.020)
F-statistic		19.158	19.420	43.700	41.970	19.158	19.420	43.700	41.970
(*p*-value)		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
J-statistic		10.521	12.167	11.093	11.695	9.201	10.401	11.174	10.855
(*p*-value)		(0.838)	(0.732)	(0.804)	(0.765)	(0.905)	(0.845)	(0.799)	(0.818)
U-statistic		29.515	51.712	16.559	16.592	30.009	2.869	17.042	11.449
(*p*-value)		(0.009)	(0.000)	(0.280)	(0.279)	(0.008)	(0.999)	(0.254)	(0.650)
Andrews test		19.441	17.985	18.688	26.150	12.076	13.856	17.246	24.904
(*p*-value)		(0.756)	(0.689)	(0.333)	(0.711)	(0.778)	(0.667)	(0.378)	(0.867)

Notes: All specifications are estimated with GMM. For further details, see Table 1.

Figure 1

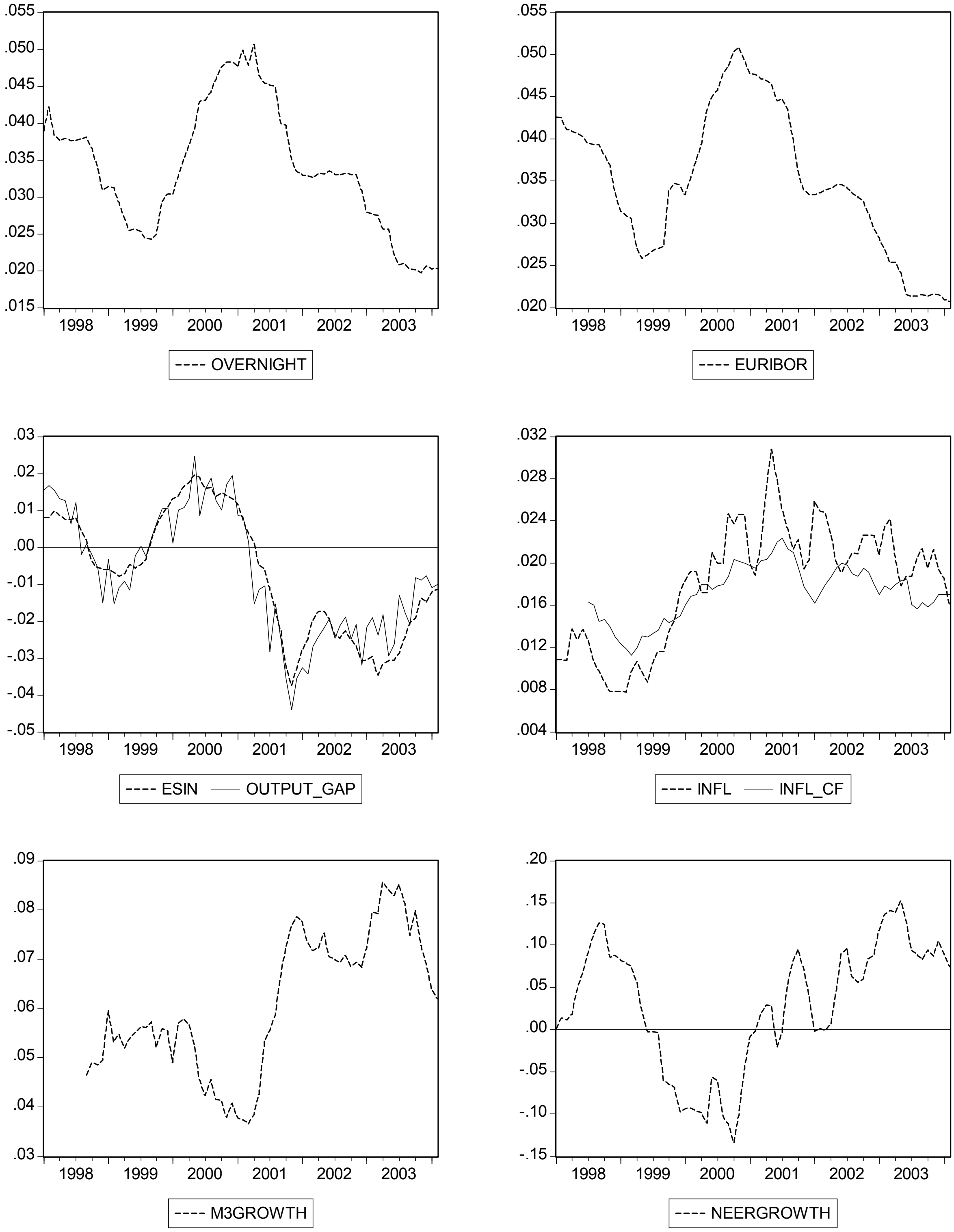
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Figure 2: The baseline model

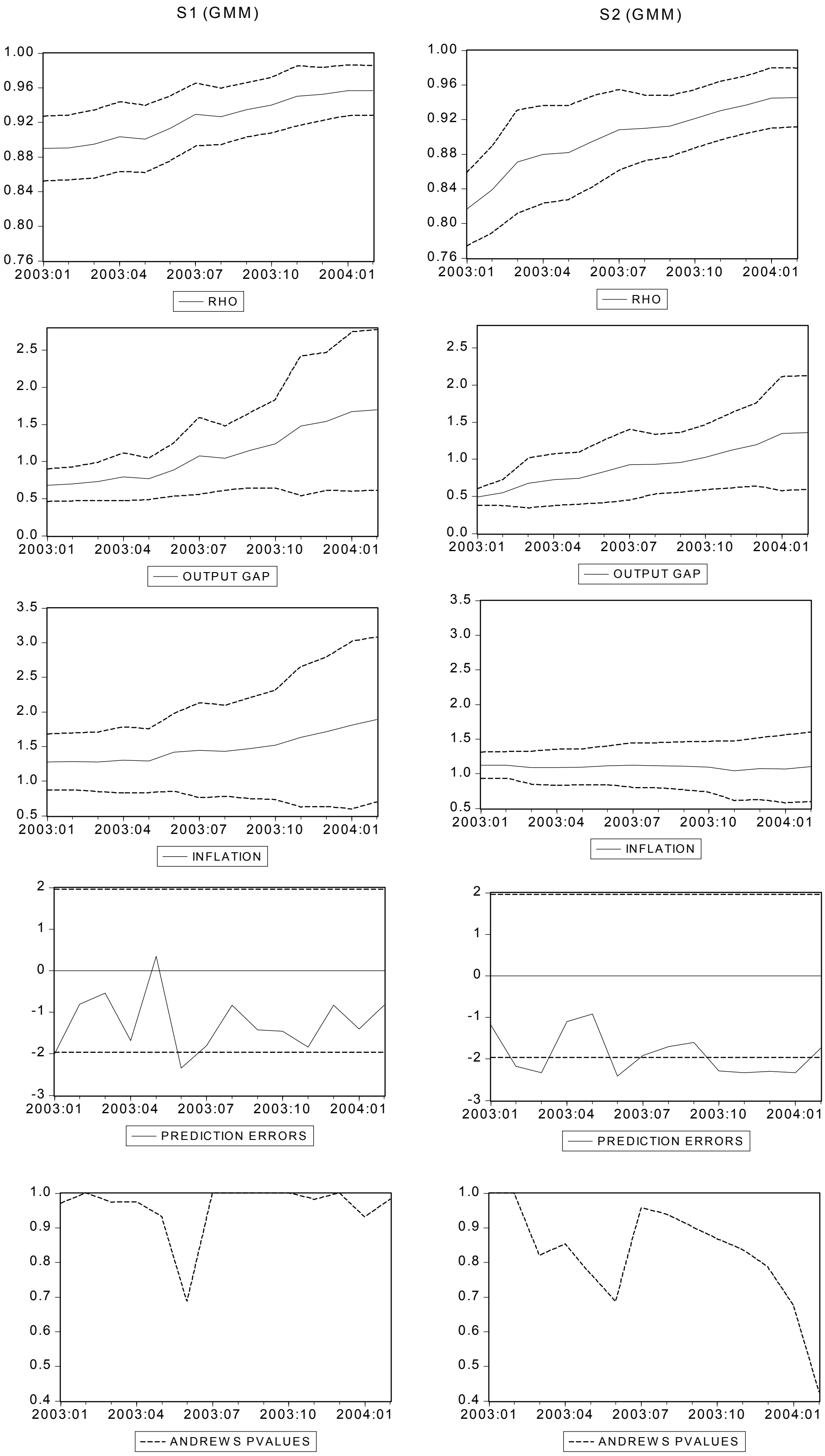

Figure 3: The baseline model

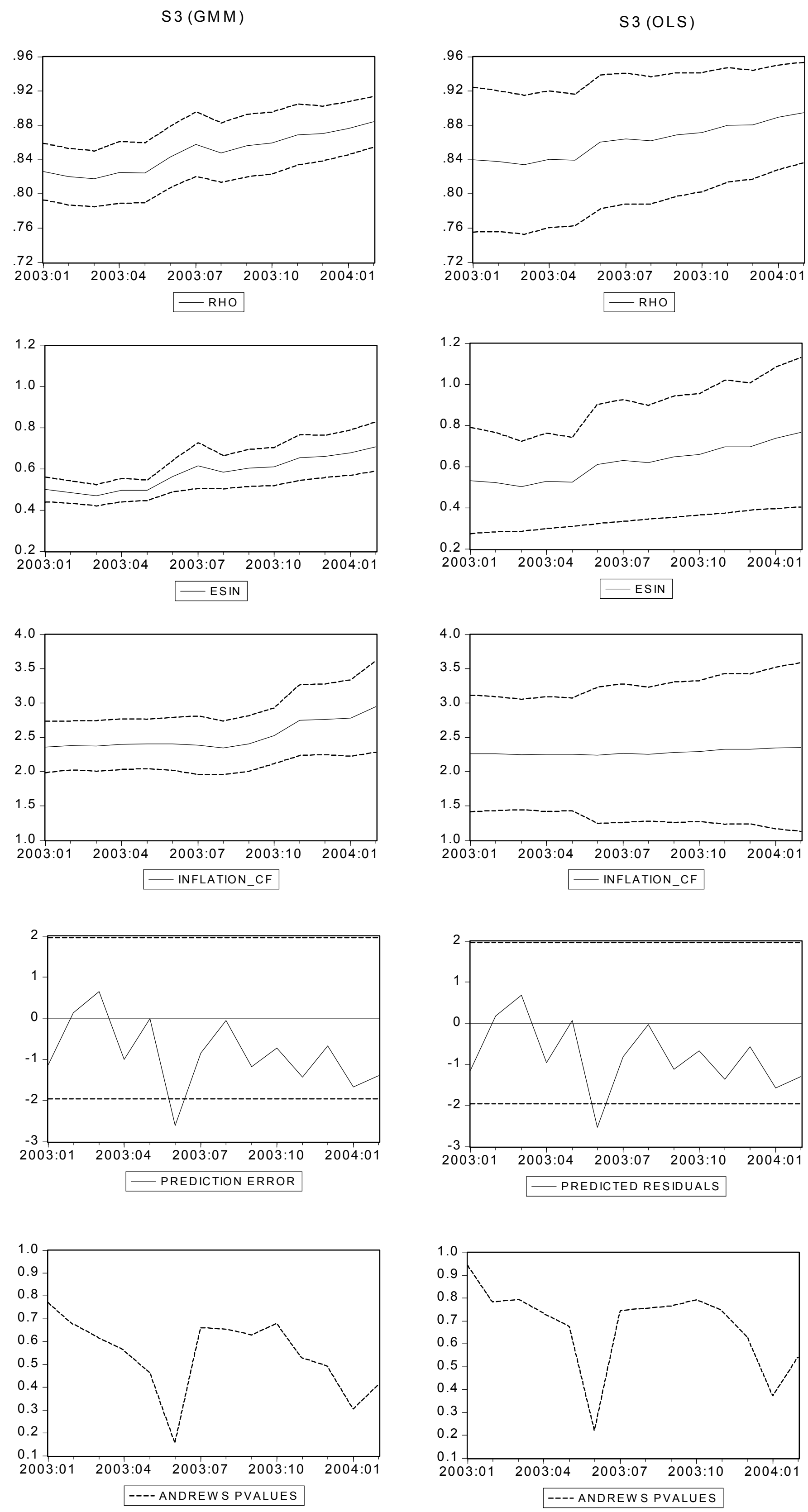

Figure 4: The baseline model augmented with money growth

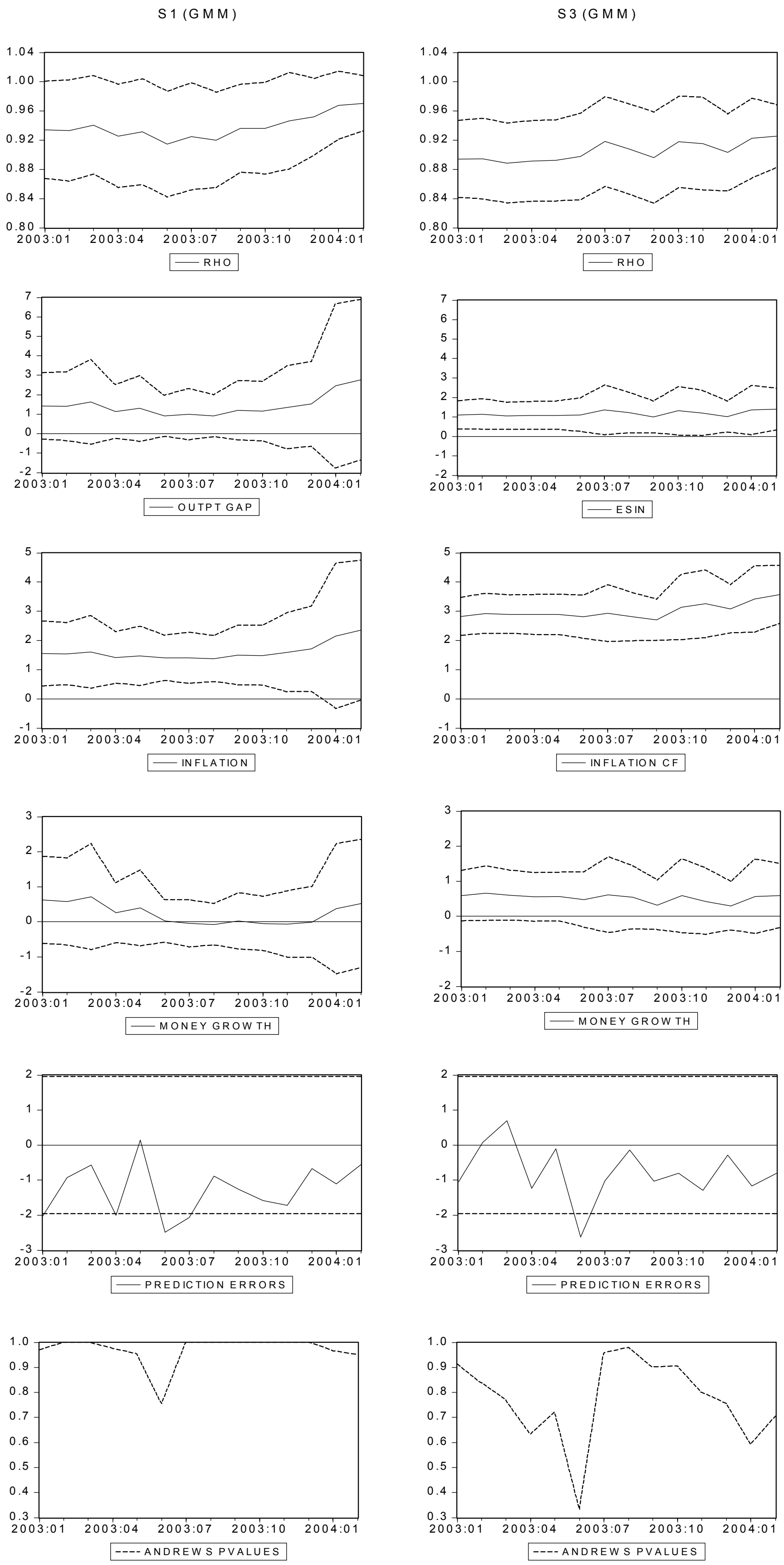

Figure 5: The baseline model augmented with exchange rate growth