Taylor Rules and Interest Rate Smoothing in the Euro Area

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Abstract

Conventional wisdom suggests that Central Banks implement monetary policy in a gradual fashion. Some researchers claim that this gradualism is due to 'optimal cautiousness'; by contrast, Rudebusch (2002) states that the observed policy-rate sluggishness is mainly due to serially correlated exogenous shocks. In this paper we employ models in first-differences to assess the 'endogenous' vs. 'exogenous' gradualism hypothesis for the Euro Area. Our results suggest that the *joint* formalization of the two hypothesis is likely to offer the best simple approximation of the Euro Area monetary policy conduct.

JEL classification system: E4, E5.

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1 Introduction

The Taylor (1993) rule has captured the attention of researchers involved in monetary policy analyses for more than a decade now. One of the reasons of its success is that, in spite of its simplicity, this rule (which links the inflation rate and a measure of the output gap to the monetary policy rate) provides a good *ex-post* description of the monetary policy implemented by various Central Banks all over the world. Interestingly, when estimating Taylor-type rules econometricians typically find that the fit of such rules remarkably improves when the *lagged* policy rate is included among the regressors. The significance and high magnitude of the lagged interest rate has stimulated several scholars in this field to investigate the rationale behind this apparent gradualism in the conduct of monetary policy, gradualism often labelled as 'interest rate smoothing', or 'monetary policy inertia'.[1]

Such a policy inertia may be rationalized in different ways. Mishkin (1999) argues that monetary authorities are very averse to reversing the policy rate course too frequently because of credibility problems, i.e. sudden, large reversals might lead agents in the economy to reduce their confidence in the Central Bank (CB henceforth)'s competence. Goodfriend (1991) discusses how a too volatile policy rate might induce financial instability because of the likely over-reaction of the markets (e.g. drastic portfolio reallocations, sharp modifications in the cost of loans for firms) that could lead to disastrous economic feedbacks. Amato and Laubach (1999) and Woodford (1999, 2001) demonstrate that a smooth policy rate path can be seen as an optimal choice when the CB is endowed with a commitment technology. In fact, an inertial rate, perceived as such by forward-looking private agents, may contribute to the reduction of the inflation bias arising under discretion. Following the intuition provided by Brainard (1967), Söderström

[1] Clarida, Galí, and Gertler (2000) estimate such a partial adjustment degree with various specifications of the Taylor rule with US data, finding a magnitude $\simeq 0.8$. Approximately the same magnitude is found by Kozicki (1999), Amato and Laubach (1999), and Domenéch, Ledo, and Taguas (2002). Estimates for some other industrialized countries are present in Henderson and McKibbin (1993) and Clarida, Galí, and Gertler (1998), while for the Euro area there exist contributions by e.g. Peersman and Smets (1999), Taylor (1999), Gerlach and Schnabel (2000), Domenéch, Ledo, and Taguas (2002), Surico (2003), Sauer and Sturm (2003), Hayo and Hofmann (2003), and Gerlach-Kristen (2003).

(1999) and Sack (2000) show that parameter uncertainty may be another element suggesting gradualism to a monetary authority whose knowledge of the monetary transmission dynamics is limited. Positive exercises conducted by Favero and Milani (2005) and Castelnuovo and Surico (2004) suggest that model uncertainty is likely to have been a very important issue for the Fed. Finally, Orphanides (2003) argues that monetary authorities respond moderately to perceived shocks because it is mindful not to respond to noise in the data.[2]

Interestingly enough, Rudebusch (2002) goes against the conventional wisdom and claims that *the interest rate smoothing at quarterly frequencies is just an illusion.* In a nutshell, his reasoning is the following. If the partial adjustment strategy had such a high importance in the policy rate setting, then rational agents should be capable to predict future values of the quarterly rate with a high degree of precision. On the contrary, standard term structure regressions show how unpredictable the policy rate is over one quarter. Rudebusch takes this evidence as convincing to claim that the *quarterly* interest rate smoothing is just negligible, and that the persistency of the observed policy rate is probably due to serially correlated *deviations* from the Taylor rate, due e.g. to commodity price scares, credit crunches, financial crises.[3]

Indeed, the issue of dynamics is important from a policy perspective. In fact, in the last two decades we have observed an improvement of the inflation-output gap trade-off in many industrialized countries. Part of this improvement is surely attributable to better monetary-policy management, as remarked by Cecchetti, Flores-Lagunes, and Krause (2006) and Favero and Rovelli (2003).[4] In general, it is necessary to understand the deter-

[2] Discussions concerning the interest rate smoothing issue may be found in Lowe and Ellis (1998), Goodhart (1999), Sack and Wieland (2000), Cecchetti (2000), and Srour (2001).

[3] Moreover, Rudebusch (2002) claims that there might be an omitted variable problem in standard Taylor rules; a similar opinion is expressed by Söderlind, Söderström, and Vredin (2005). Indeed, if the Taylor model is misspecified and missing a important serially correlated regressors, then the importance of the lagged interest rate might be just spurious. We deal with this relevant issue later in the paper.

[4] Both Cecchetti et al (2001) and Favero and Rovelli (2003) acknowledge that the improved inflation-output gap trade-off has probably not been uniquely caused by a better

minants of this successful management, in order to possibly replicate this success in presence of similar macroeconomic conditions. Then, is the observed gradualism *endogenous,* i.e. stemming from the systematic component of the monetary policy under analysis, or *exogenous*, i.e. due to serially correlated policy shocks?

While the discussion has been quite lively as far as the U.S. case is concerned [Rudebusch (2002), English, Nelson, and Sack (2003, ENS hereafter), Castelnuovo (2003a)], to our knowledge the literature is still silent with regard to the Euro Area case. In fact, although several contributions about the 'counterfactual' as well as the true European Central Bank have already focussed their attention on Taylor-type rules,[5] none of them has deepened the important issue of dynamics discussed above.

In this paper we test the null hypothesis of absence of any interest rate smoothing mechanism in the Euro Area context. In doing so, we take into account several definitions of the Taylor rate, in order to control for possible omitted variables problems as done by Clarida, Galí, and Gertler (1998, CGG henceforth), Gerlach and Schnabel (2000), Surico (2003), and Gerdesmeier and Roffia (2004a). According to the simple test we implement, there is evidence in favor of a significant impact exerted by the partial adjustment mechanism on the dynamics of the policy rate. Importantly, this does not imply that the serial correlation of the policy rate is excluded from the picture. In fact, both mechanisms are likely to play a role in shaping the path of the policy rate in the Euro Area.

In performing this exercise, we have to keep in mind some important caveats. First, this is an *ex-post* analysis mostly referring to a *counterfactual* monetary policy conduct. In fact, the European Central Bank began to manage the Euro Area monetary policy in 1999, while the sample we employ starts much earlier, i.e. at the beginning of the year 1980. Second, we deal with a dataset created by computing weighted-averages for the relevant data

monetary policy management. In fact, there is evidence of a change in monetary policy preferences, and of more favourable sequences of supply shocks. Still, better monetary policy management seems to have been quite significant for the last two decades now. For a contribution focussing on the measurement of policy-makers' preferences over inflation and the output gap volatilities, see Cecchetti et al (2002).

[5] A nice survey of such contributions is offered by Sauer and Sturm (2003).

for different, potentially 'heterogeneous' countries such as those belonging to the Euro Area. Third, we deal with revised-data, while a CB operates in real-time.[6]

Nevertheless, although some breaks may be clearly identified in this pattern (e.g. ERM crisis in 1992), a common effort to bring down inflation to more sustainable levels has been implemented by several European countries since the early '80s, and continued under the monetary policy management by the European Central Bank [Gerdesmeier and Roffia (2004a)]. Moreover, the use of synthetic European data is fairly widespread among researchers [e.g. Peersman and Smets (1999), Taylor (1999), Gerlach and Schnabel (2000), Doménech et al (2002), Gerdesmeier and Roffia (2004a,b), Surico (2003), Sauer and Sturm (2003)]. Therefore, we think that our exercise can be considered as a fairly good first approximation of the track followed by the 'average' monetary policy conduct in the Euro Area during the last two decades. Given what written above we obviously refrain from attaching any normative evaluation to our estimated simple Taylor rules.[7]

The structure of the paper reads as follows. Section 2 explains the advantage of employing models in first differences when dealing with the identification issue affecting models in levels. In Section 3 we present and discuss our findings. Section 4 concludes. A description of the dataset employed in this paper is offered to the reader, and References follow.

2 A direct test for partial adjustment versus serial correlation

Thinking of models in levels, an econometrician can easily build up two frameworks for representing the partial adjustment (PA) vs. the serial correlation (SC) hypothesis. In particular, the former may be captured by the

[6] Gerdesmeier and Roffia (2004b) show that Orphanides (2001)'s intuition on the importance of dealing with the real-time data issue applies to the Euro Area as well. By contrast, Sauer and Sturm (2003) demonstrate that the use of real-time industrial production data does not seem to play a very significant role for the point estimates of the Taylor rules they focus on. We leave the assessment of the impact of the data-revision issue on the results presented in this study to future research.

[7] For a critical assessment of Taylor-type rules in such a context, see European Central Bank (2001).

following model:

$$i_t = (1-\rho)\widetilde{i}_t + \rho i_{t-1} + \eta_t \tag{1}$$

where i_t is the short-term policy rate managed by the CB in order to influence the inflation rate and the business-cycle, $\widetilde{i}_t$ is the target rate (i.e. Taylor rate), ρ measures the importance of the interest rate smoothing motive, and η_t is a white noise policy shock. Alternatively, a process relating the serially correlated policy shock to the policy rate with no-endogenous persistence may be shaped as follows:

$$i_t = \widetilde{i}_t + \varepsilon_t,\ \varepsilon_t = \rho_\varepsilon \varepsilon_{t-1} + \eta_t \tag{2}$$

where ε_t is an AR(1) process with root ρ_ε.[8]

Following ENS (2003) and Castelnuovo (2003a,b), we manipulate the 'PA' model (1) and the 'SC' mechanism (2) in order to work with their first-difference counterparts. Once done so, we are left with the following equations for the PA vs. SC hypothesis:

$$\Delta i_t = (1-\rho)\Delta\widetilde{i}_t + (1-\rho)(\widetilde{i}_{t-1} - i_{t-1}) + \eta_t \tag{3}$$

vs.

$$\Delta i_t = \Delta\widetilde{i}_t + (1-\rho_\varepsilon)(\widetilde{i}_{t-1} - i_{t-1}) + \eta_t \tag{4}$$

The latter equation sheds some light on the implications of the SC engine. Here, variations of the Taylor-rate cause an immediate and full reaction of the policy rate change. In fact, there is no inertial adjustment, which is by contrast present in equation (3) via the coefficient $(1-\rho)$. Then, it is possible to build up a direct test on the PA vs. SC hypotheses by exploiting the empirical model

[8] We performed some econometric exercises in order to measure which is the serial correlation order featuring the residuals of simple backward and forward looking Taylor rules without smoothing. Our findings suggest that an AR(1) process is a good approximation of the errors. These findings - not included in the paper for sake of brevity - are available upon request.

$$\Delta i_t = \gamma_2 \Delta \widetilde{i}_t + \gamma_3 (\widetilde{i}_{t-1} - i_{t-1}) + \eta_t \tag{5}$$

and testing the null hypothesis

$$H0_{SC} : \gamma_2 = 1 \tag{6}$$

Under the null (6), the SC specification holds true. By contrast, a rejection of the null hypothesis has clear implications for the dynamics of the policy rate, which must be at that point influenced *also* by its lag. Clearly, this does not imply that the pure PA model is the only alternative to the null. More likely, the two mechanisms jointly shape the policy rate path. The point we aim at making here is that the rejection of the null (6) implies that the lagged policy rate enters the Taylor-type rule in its own right.

Taylor-rate definitions employed in our study

As far as the Taylor rate $\widetilde{i}_t$ is concerned, it is natural to concentrate on some popular definitions of it. Our benchmark is the original Taylor (1993) rate, which reads as follows:

$$\widetilde{i}_t = c + b_\pi \pi_t^{HICP} + b_y y_t \tag{7}$$

where c is a constant, π_t^{HICP}= year-on-year HICP inflation rate, and y_t = the output gap.[9],[10] A different specification of the Taylor rate has been popularized by CGG (1998, 2000). These authors have underlined the importance for the CB to adjust the policy rate with respect to *future*, forecast movements of both inflation and output gap. Their idea finds its rationale

[9] For a description of the dataset employed in our study, as well as the construction of the variables involved in our regressions, see the Data description at the end of the paper. The dataset we used is available upon request.

[10] In Taylor (1993), the policy rule reads as follows: $i_t = \pi_t + 0.5 y_t + 0.5(\pi_t - \pi^*) + r^*$, with $\pi^* = r^* = 2\%$. Then, the constant c in the various Taylor rates is a linear convolution of the inflation target π^* and the real interest rate of equilibrium r^*, i.e. $r^* - b_\pi \pi^*$. Neither in Rudebusch (2002)'s nor in our study the focus is the one of assessing these elements; for investigations concentrating on these components, see Judd and Rudebusch (1998) and Domenéch, Ledo, and Taguas (2002).

in the lags affecting the monetary policy transmission. Their definition of the Taylor rate can be captured by the following modelization:[11]

$$\widetilde{i}_t = c + b_\pi E_{t-1}\pi_{t+4}^{HICP} + b_y E_{t-1} y_t \tag{8}$$

However, as already mentioned above, Rudebusch (2002) calls for omitted serially correlated variables as potential cause of the estimated high degree of PA. To check also for this, we enrich the original specification (7) by adding a third regressor, as follows:

$$\widetilde{i}_t = c + b_\pi \pi_t^{HICP} + b_y y_t + b_z z_t \tag{9}$$

In our exercise, the regressor z_t plays different roles. A variable that we want to control for is a quadratic transformation of the output gap level, i.e. $z_t = y_t^2$. In doing so we feel inspired by recent works on CBs' asymmetric preferences, which imply a non-quadratic representation of their loss function.[12] Many normative analyses conducted so far have relied on a quadratic formalization of the CB's penalty function. Indeed, apart from analytical tractability, there does not seem to be an obvious reason why a CB should symmetrically target the output gap measure [Blinder (1997), Goodhart (1999), Mayer (2002)]. With our simple modeling strategy we try to capture possibly asymmetric reactions by the CB to business-cycle movements.

Moreover, we also aim at investigating the CB's possible responses to movements in variables such as money ($M3$) and the nominal effective exchange rate, on the lines of contributions such as CGG (1998), Gerlach and Schnabel (2000), and Gerdesmeier and Roffia (2004a). The first element was important for the Bundesbank (CGG, 1998) and it has still a prominent status within the ECB's monetary policy strategy [European Central Bank (2003)]; by contrast, the latter component is meant to capture possible 'external pressures' affecting Euroland. Given that a CB reacts to deviations

[11] Sauer and Sturm (2003) underline the importance of considering a forward-looking Taylor rule when describing the monetary policy implemented in the Euro Area.

[12] Researchers such as Gerlach (2003), Surico (2002, 2003), and Cukierman and Muscatelli (2003) have performed empirical endeavours along this avenue. See also the references quoted in those papers.

of the relevant aggregates from their long-run equilibrium values or reference values, we estimate our policy rules by taking the nominal effective exchange rate in deviations with respect to its sample mean. As far as the $M3$ growth rate is concerned, we consider either its deviations with respect to the constant reference value 4.5%, or with respect to the time-varying reference value computed by Gerlach and Svensson (2003).[13] Finally, as an alternative to the use of future realized inflation in eq. (8), we also employ survey data on one-year ahead inflation expectations provided by Consensus Economics.[14]

We consider the sample 1980Q1-2003Q4. We adopt a Nonlinear Least Square estimator for models without expectations (i.e. when either (7) or (9) is considered) and when survey data are employed. By constrast, we employ a 2-Stage Nonlinear Least Square procedure when (8) is taken into account and agents are assumed to be fully rational.[15]

In the next Section we present and comment our findings.

3 Findings

The test for the Euro area rejects the *pure* SC mechanisms hypothesis formalized in eq. (2). This is understandable when looking at Table 1, that displays the results stemming from the implementation of the ENS test. Notably, the null (6) is strongly rejected with all the different specifications of the Taylor rate considered. This implies that the PA mechanism played a role in explaining the dynamics of the European short-term policy rate. As

[13] The construction of the nominal money growth rate gap, i.e. $\Delta_4 m_t - \Delta_4 m_t^*$, is detailed in the Data description. Notice that the reference value is set equal to $4\frac{1}{2}\%$ since 1999Q1 (included).

[14] Consensus Economics is the world's leading international economic survey organization and polls more than 600 economists each month to obtain their forecasts and views.

[15] The initial conditions for the NLS/2SNLS are provided by LS/2SLS regressions. The instruments used for our 2SLS regressions are a constant and 5 lags of the HICP inflation rate, of the output gap, and of the short-term nominal interest rate. Such number of lags was selected by running an unrestricted VAR(n) with HICP inflation and the output gap as endogenous variables, and the policy rate (lags from 1 to n) as exogenous variable, and by checking the indications stemming from standard lag-length criteria. Likelihood-ratio, Final prediction error, Akaike, and Hannan-Quinn criteria all suggested a number of lags equal to 5.

previously pointed out, this does not necessarily rule out the importance of the SC mechanism in triggering interest rate movements in the Euro Area. Most likely, both mechanisms played an important role in the sample under consideration.

Interestingly enough, almost all the point estimates of the inflation coefficient b_π suggest that a fairly tight monetary policy was implemented in Europe in the '80s and '90s.[16],[17] Indeed, all these simple feedback rules find in the output gap measure a significant regressor, so confirming also for the Euro Area the goodness of Taylor (1993)'s descriptive scheme.[18] By contrast, the additional regressors considered here do not show any statistical relevance at the standard confidence levels. This result is in line with those found by Peersman and Smets (1999) and Gerlach and Schnabel (2000).[19]

[16] According to the standard New-Keynesian model a la Clarida et al (1999), the necessary and sufficient condition for having a unique and stable equilibrium in an economic system populated by rational agents is (approximately) $b_\pi > 1$. Interestingly, with the estimates at hand we can never statistically reject the null hypothesis of unique and stable equilibrium.

[17] To assess the strenght of the instruments employed for estimating the model with forward looking agents (when rational expectations were imposed), we computed the Cragg-Donald statistic [minimum eigenvalue of the matrix analog of the F-statistic from the first stage regression of TSLS, see Stock and Yogo (2004)]. The computed value amounts to 0.7667, lower than the critical values proposed for rejecting the null of weak instruments at an acceptable level. Some single reduced form regressions confirmed that this result is largely due to the difficulty of instrumenting the 4-quarters ahead first difference of the inflation rate. Other sets of instruments did not improve the result. However, the J-test confirms the orthogonality of the instruments to the error in the regression (p-value: 0.287890). Moreover, the regression with survey data confirmed the descriptive power of the forward looking model.

[18] The figures displayed in Table 1 refer to Taylor rules estimated with an output gap measured as log-deviation of the real GDP with respect to a linear trend (with constant), i.e. our benchmark case. Our results in terms of statistical importance of the output gap in such regressions turns out to be robust when an HP-filter measure of the potential output is employed, as well as when the output gap provided in the Area-Wide Model database is taken into account. The figures of the latter two cases are available upon request.

[19] Instead, these findings seem to be at odds with those in CGG (1998), whose investigated sample spans from 1979 up to 1993. One reason for this different findings may rely on the different data at hand: National in CGG (1998)'s case, aggregate in ours. Moreover, a plausible explanation for these contrasting results may be the fact that the Maastricht Treaty, signed up in 1992, forced all the signatory countries to implement tight monetary and fiscal policies in order to quickly converge toward the Maastricht criteria. Then, although important, external pressures might have been replaced by domestic concerns fully captured by our sample choice, while only partially by CGG's.

[insert Table 1 about here]

Stability analysis

Table 2 shows the outcome of a stability analysis performed by exploiting two very popular stability tests: The Chow-breakpoint test and the Chow-forecast test.[20] As a break-date we chose 1999Q1, i.e. the beginning of Stage Three of EMU, a very important date for the Euro Area from an economic perspective. Overall, our estimates turn out to be fairly stable. In particular, the only doubts regarding the stability of our coefficients are cast on the Forward-looking model (both when the assumption of rational expectations is explicitely imposed and when we employ survey data) and on the M3 growth rate model with the constant reference value. Nevertheless, the rejection of the stability of the estimated coefficients is far from being overwhelming, given the different and conflicting indications coming from the two tests we employed.

[insert Table 2 about here]

Variance decomposition analysis

To gauge the relative importance of the interest rate smoothing mechanism vs. the serial correlation model we compute the variance decomposition of the deviations of the actual rate from the desired Taylor rate, i.e. the 'interest rate gap' $i_t - \widetilde{i}_t$. Consider the following model:

$$\begin{aligned} i_t &= (1-\rho)\widetilde{i}_t + \rho i_{t-1} + \varepsilon_t \qquad (10) \\ \varepsilon_t &= \rho_\varepsilon \varepsilon_{t-1} + \eta_t \end{aligned}$$

[20] The idea of the Chow-breakpoint test is to fit the equation at hand separately for each subsample to see whether there are significant differences in the estimated equations. A significant difference indicates a structural change in the relationship. Differently, the Chow-forecast test first estimates the model for a subsample comprised of the first observations, then exploits such estimated model to predict the values of the dependent variable in the remaining data points. A large difference between the actual and predicted values casts doubt on the stability of the estimated relation over the two subsamples. For more information about these tests, see Greene (1997), chapter 7.

After some manipulations, it is possible to derive a 'hybrid' formula describing the path of the interest rate gap, i.e.

$$i_t - \widetilde{i}_t = \rho(i_{t-1} - \widetilde{i}_t) + \rho_\varepsilon \varepsilon_{t-1} + \eta_t \tag{11}$$

We concentrate on the standard specification of the Taylor rate (7) with $b_z = 0$, and compute $\widetilde{i}_t$ by exploiting the point estimates displayed in Table 1 and the actual time series of inflation and the output gap.[21] Our least-squares estimates of eq. (11) read as follows:[22]

$$i_t - \widetilde{i}_t = \underset{[0.04]}{0.84}(i_{t-1} - \widetilde{i}_t) + \underset{[0.10]}{0.44}\varepsilon_{t-1} + \widehat{\eta}_t$$

Then, we perform the following Montecarlo exercise. We draw $10,000$ series of 96 values (96 being the sample-size at hand) of η from the empirical distribution of the estimated residuals. For each series of residuals, we employ the estimated parameters of eq. (11) and the time series of the Taylor rate $\widetilde{i}_t$ to simulate time-series of the policy rate i_t and the serially correlated error term ε. Then we re-estimate the model (11) for each of the $10,000$ samples, and calculate the variance decomposition for each sample.[23]

Figure 1 plots the distributions of the estimated parameters, along with those of the share of the explained variance. As it is possible to see, there is uncertainty both on the estimated values and on the contribution that the two mechanisms under investigation bring to explain the volatility of the interest rate gap.[24] The plotted distributions seem to point towards a

[21] In particular, $\widetilde{i}_t = \widehat{c} + \widehat{b_\pi}\pi_t^{HICP} + \widehat{b_y}y_t = 2.54 + 1.15\pi_t^{HICP} + 0.98y_t$, with y_t being the linearly-detrended real GDP in logs.

[22] Sample: 1980Q1-2003Q4. $\overline{R}^2 = 0.98, \sigma_{i-\widetilde{i}} = 3.62, \widehat{\sigma_\eta} = 0.44$. A spike dummy for the observation 1992Q3 turned out to be highly significant ($p - value = 0.00$), with an estimated coefficient equal to 1.18.

[23] Of course, the variance of the interest rate gap $i_t - \widetilde{i}_t$ is explained by the variance of $i_{t-1} - \widetilde{i}_t$, that of the autoregressive element ε_{t-1}, that of the white noise error η_t, and the covariances existing between these elements. In our analysis we concentrate on the variance explained by the PA and SC mechanisms. To correctly identify the contribution of the former to the total explained variance (contribution which is actually boosted by the presence of the serially correlated error term), we compute it by setting $\rho_\varepsilon = 0$ as done by ENS (2003).

[24] As in ENS (2003), the mean of the distribution of the ρ parameter turns out to be lower than that of the 'true' value, due to the small-sample bias affecting the least-square estimator in this context.

somewhat higher relative importance of the PA mechanism in explaining the interest rate gap volatility, although the exact quantification of the share of explained volatility by PA is difficult due to the uncertainty surrounding it.[25] However, the participation of both these processes to the formation of the interest rate gap turns out to be quite supported by the data, with both the 95% confidence intervals excluding the zero-value.

[**insert Figure 1 about here**]

4 Conclusions

In this paper we focussed our attention on the interest rate smoothing argument in Taylor-type schemes. Following English, Nelson, and Sack (2003) and Castelnuovo (2003a,b), we implemented a direct test for the interest rate smoothing hypothesis in the Euro Area case. Our evidence supports the empirical relevance of both endogenous and exogenous persistence when fitting simple Taylor rules for the Euro Area. As far as the point estimates we obtained are concerned, our evidence stresses the tight reaction implemented in the Euro Area to inflation fluctuations, as well as the role played by business cycle fluctuations in influencing the policy rate.

Admittedly, our exercise missed to properly consider the role possibly played by expectations on future monetary policy actions embedded in the term structure of interest rates [see Rudebusch (2005) and the literature cited therein]. An extension of this study along such avenue is in our agenda.

5 Data description

For our study we employed the *Area Wide Model* dataset. This dataset collects seasonally adjusted series (except for the HICP index) built up with the 'Index aggregation method', i.e. the log-level index for any series is constructed as a weighted average of the log-level country-specific indexes

[25] Interestingly, the 95% confidence interval for the share explained by PA spans from 74% to 92%, i.e. it is much tighter than the one found by ENS (2003) in the American case (i.e. 7% to 91%). This might be due to the much larger sample size available, i.e. 96 observations in our study vs. 56 in theirs.

of the 12 countries belonging to the Euro Area. A standard adjustment for seasonality - i.e. the 'Ratio to moving average (multiplicative)' was applied to the HICP index, and it revealed there was no need of performing any seasonal-adjustment for such series. The series employed in this studies are the following (labels as in Fagan et al, 2005): 'HICP'= harmonized index of consumer prices, 'YER' = real GDP, 'STN' = short-term nominal interest rate, 'EEN' = effective nominal exchange rate, 'YGA' = the output gap. The year-on-year HICP inflation rate was computed as $\pi_t^{HICP} = 100[\log(HICP_t/HICP_{t-4})]$. Our benchmark measure of the output gap was computed by applying a linear trend (with constant) to the log of real GDP. As alternatives, we computed the potential output by employing i) either a standard HP-filter approach with smoothing parameter = 1600, or ii) the output gap measured by Fagan et al (2005), i.e. by considering a potential output costructed with the production function approach. The nominal exchange rate used in our regressions was demeaned by considering the in-sample mean for 1980Q1-2003Q4. For more information on the above mentioned time-series, see Fagan et al (2005).

The measure of inflation forecasts employed is the one produced by Consensus Economics/Consensus Forecast. To construct the forecast for the Euro Area, we pooled the 1-year ahead forecasts on consumer price index (percentualized changes over the previous year) by weighting them as indicated in the *Orange Book* (December 2002, Table 8.1, page 184, definition: Real Consumer Spending Weights, PPP exchange rates). We pooled data regarding Germany (weight: 0.3341), France (0.2380), Italy (0.2333), Spain (0.1278), and the Netherlands (0.0668). The transformation from monthly to quarterly frequencies was made by taking the within-quarter average values.

The annual growth rate of the money stock '$M3$' was downloaded from the European Central Bank's web-site (http://www.ecb.int, under 'Key indicators'), and it refers to a seasonally-adjusted measure of $M3$. We performed a 'monthly-to-quarterly' frequency transformation by taking simple averages of monthly observations. The $M3$ growth rate employed in our regressions was considered in deviations with respect to the 4.5% reference value.

The construction of the nominal money growth rate gap, i.e. $\Delta_4 m_t - \Delta_4 m_t^*$, requires the specification of the time-varying reference value $\Delta_4 m_t^*$. According to the quantity theory of money, $\Delta_4 m_t^* \equiv \Delta_4 \widehat{p}_t + \Delta_4 y_t^* - \Delta_4 v_t^*$, where $\Delta_4 \widehat{p}_t$ stands for the average inflation objective for the Euro Area countries, $\Delta_4 y_t^*$ is the yearly growth rate of the potential output, and $\Delta_4 v_t^*$ is the yearly growth rate of velocity. Following Gerlach and Svensson (2003), $\Delta_4 \widehat{p}_{t+1} = \Delta_4 \widehat{p}_{t+1}^b + 0.964(\Delta_4 \widehat{p}_t - \Delta_4 \widehat{p}_t^b)$, with $\Delta_4 \widehat{p}_t^b$ being the Bundesbank's inflation objective [source: Gerlach and Svensson (2003), and set to 1.5% from 2001Q2 onwards]. The potential output was approximated by linearly detrending the real GDP in logs (sample: 1970Q1-2003Q4). The measure of long-run velocity was defined as $v_t^* \equiv (1 - 0.98)y_t^* - 0.31 timetrend + 0.85(\overline{i_{long}} - \overline{i})$, $\overline{i_{long}} = 8.83$ and $\overline{i} = 8.07$ being the full-sample averages of - respectively - the long-term interest rate and the policy rate. The estimated values employed for constructing the series were taken from Tables 1 and 2 in Gerlach and Svensson (2003).

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Taylor rate specification	b_π	b_y	b_z	γ_2	γ_3	$\overline{R}^2$	H_0:γ_2=1 ($p-value$)
Standard specification	1.15*** [0.28]	0.98*** [0.22]	–	0.34*** [0.10]	0.09*** [0.03]	0.27	0.00***
Forward looking	1.00* [0.60]	1.80** [0.85]	–	0.13 [0.14]	0.07** [0.03]	0.18	0.00***
F. look./Consens. Forec.	2.00*** [0.21]	0.59*** [0.13]	–	0.30** [0.11]	0.39*** [0.06]	0.64	0.00***
Asymmetric preferences	1.15*** [0.28]	0.98*** [0.23]	0.00 [0.06]	0.34*** [0.10]	0.09*** [0.03]	0.26	0.00***
Nominal effect. exch. rate	1.25*** [0.23]	0.95*** [0.20]	1.09 [3.14]	0.40*** [0.10]	0.10*** [0.03]	0.30	0.00***
M3 growth rate - 4.5%	1.19*** [0.22]	0.89*** [0.20]	0.18 [0.22]	0.43*** [0.11]	0.10*** [0.03]	0.31	0.00***
M3 growth rate gap	1.31*** [0.22]	0.96*** [0.17]	−0.03 [0.18]	0.39*** [0.10]	0.12*** [0.05]	0.32	0.00***

Estimated model: Eq. (5) in the text (a constant and a spike dummy for 1992Q3 also included).
Point estimates [Newey-West corrected standard errors in squared brackets].
*=90%/**=95%/***=99% statistical confidence for rejecting the null H_0: insign. coefficient.
H_0:γ_2 tested with a standard Wald test. Estimates performed via NLS for the backward-looking rule; 2SLS for the forward-looking one (instruments empl.: $[c,\pi_{t-1}^{HICP}..\pi_{t-5}^{HICP}, y_{t-1}..y_{t-5}, i_{t-1}..i_{t-5}]$).
$Analog - Ftest$: $g_{min} = 0.7667$ [evidence of weak instruments, see Stock and Yogo (2004)].
$J - stat.$: 0.142214; $J - test(p - value)$: 0.287890 [null of orthogonality: not rejected]
Sub-samples:'1'=1980Q1-1998Q4,'2'=1999Q1-2003Q4 (1st, 2nd, 4th, 5th, and 6th specification); '1'=1990Q1-1998Q4,'2'=1999Q1-2003Q1 (3rd);'1'= 1981Q2-1998Q4,'2'=1999Q1-2003Q4 (7th).

Table 1: Endogenous (PA) vs. Exogenous (SC) persistence: Test based on Models in First-Differences.

Taylor rate specification	*Chow − breakpoint* ($F-test$,$p-value$)	*Chow − forecast* ($F-test$,$p-value$)
Standard specification	0.50	0.99
Forward looking	0.03**	0.55
F. look./Consensus Forecast	0.03**	0.30
Asymmetric preferences	0.51	0.99
Nominal effect. exch. rate	0.23	0.89
M3 growth rate - 4.5%	0.08*	0.79
M3 growth rate gap	0.12	0.85

*=90%/**=95%/***=99% stat. confid. for the rej. of H_0: estimated parameters' stability.
Breakpoint: 1999Q1. Distributions of the tests ('1'=1st sub-sample,'2'= 2nd,'12'='1'+'2'):
Chow breakpoint test, F-stat: $F(k, T-2k) = \frac{[u'_{12}u_{12} - (u'_1 u_1 + u'_2 u_2)]/k}{(u'_1 u_1 + u'_2 u_2)/(T-2k)}$
Chow forecast test, F-stat: $F(T_2, T_1 - k) = \frac{(u'_{12}u_{12} - u'_1 u_1)/T_2}{u'_1 u_1/(T_1 - k)}$
Sub-samples:'1'=1980Q1-1998Q4,'2'=1999Q1-2003Q4 (1st, 2nd, 4th, 5th, and 6th specific.);
'1'=1990Q1-1998Q4,'2'=1999Q1-2003Q1 (3rd);'1'= 1981Q2-1998Q4,'2'=1999Q1-2003Q4 (7th).

Table 2: Estimated Models in First-Differences: Stability tests.

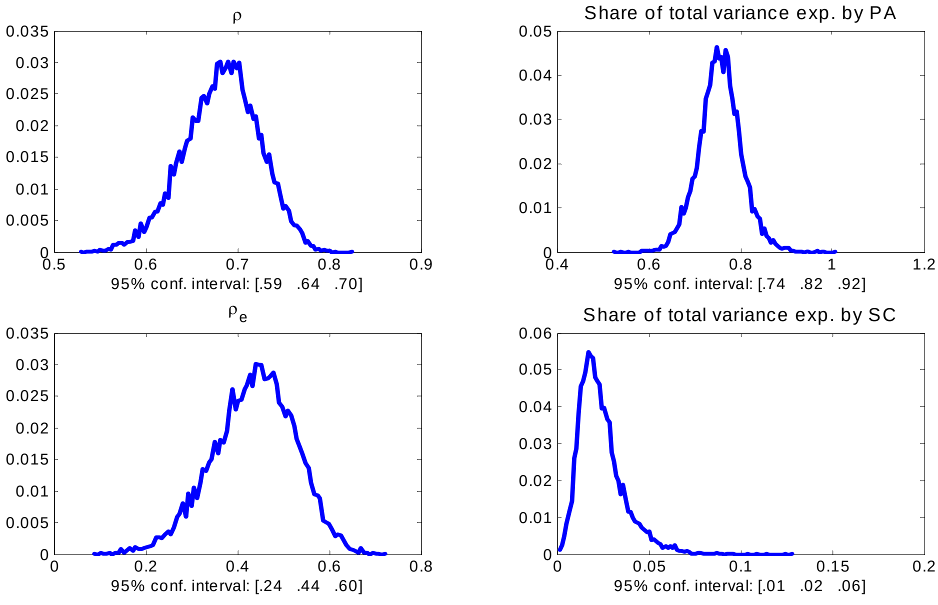

Figure 1: Montecarlo distributions of the estimated parameters and implied shares of explained total variance of the interest rate gap.