

# State Asymmetries in the Effects of Monetary-Policy Shocks on Output: Some New Evidence for the Euro-area

J.J.Dolado<sup>†</sup> and Ramón María-Dolores<sup>††</sup>,

(<sup>†</sup>)Universidad Carlos III and CEPR

(<sup>††</sup>) Universidad de Murcia

## Abstract

In this chapter, we provide some empirical evidence on whether the effects monetary policy shocks on real output growth in the Euro area depend upon the phase of the business cycle that the economy was undergoing (the so-called *state* asymmetries). To do so, we follow a two-step procedure: (i) first, we derive short-term interest rate shocks from a Taylor rule which accounts for a nonlinearity in the interest-rate setting behaviour of the central bank, and (ii) next, we apply a multivariate version of Hamilton(1989)'s Markov switching methodology to allow for different effects of interest-rate shocks on real output growth in periods of high and low growth. Our findings provide some support for the presence of this type of asymmetries, whereby interest rate shocks have larger effects in recessions than in expansions.

# 1 Introduction

Our main goal in this chapter is to provide empirical evidence for the Euro-area (EA, hereafter) on whether monetary policy shocks have had asymmetric effects on real output growth depending on the phase of the business cycle that the economy was undergoing during the period 1996-2003. More precisely, our objective is to test whether these effects are significantly different in expansions and recessions. For this purpose, we follow the methodology advocated by Garcia and Schaller (1995), Ravn and Sola (1996) and Dolado and Maria-Dolores (2001) where the well-known Hamilton's (1989) approach to model univariate processes subject to stochastic regime shifts is extended to a Multivariate Markov-Switching (MMS) framework.<sup>1</sup> In this setup, real output growth (directly) and the transition probabilities between cyclical phases, are allowed to depend on shocks to a monetary-policy (*Taylor*) rule describing the evolution of a short-term interest rate controlled by the monetary authorities. Either the coefficients on these shocks or the transition probabilities are themselves functions of the latent variable capturing regime changes, providing in this way a flexible modelling framework where to implement the above-mentioned test. The use of the MMS methodology is appropriate to analyze the cyclical effects of changes in the monetary-policy stance in the EA since, unlike what happens with the NBER dating for the US cycle, an official dating for the EA cycle is not yet officially available.<sup>2</sup> Hence, the MMS approach will enable us to address a number of interesting issues ranging from *Do monetary policy shocks have different effects on output depending on the phase in which the change in monetary policy took place?* to *Do changes in the monetary policy stance alter the transition probabilities from a recession to a boom and conversely?*

To measure the stance of monetary policy in the EA, a Taylor rule has been estimated using monthly data for the period 1996(01)- 2003(12). The choice of the sample size is dictated by the adoption of inflation targets by central banks in most of the EA countries since the mid-1990s, following the

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<sup>1</sup>Another related paper which, however, uses a different methodology (*Smooth Transition Regression, STR, models*) is Osborn *et al.* (2002) where the effects of monetary policy on output in the UK are examined.

<sup>2</sup>However, non-official dating of the EA business cycle can be found in Artis *et al.* (2004) and in the EU COIN indicator of the *Centre for Economic Performance* (CEPR) whose methodology is explained in Forni *et al.* (2005).

collapse of the EMS in September 1992. Since the European Central Bank (ECB) started its control of monetary policy in the EA in 1999, our implicit operating assumption is that the national central banks conducted monetary policy during the preparatory stages to the foundation of the ECB (under the supervision of the European Monetary Institute created in 1994), with the same principles that the ECB uses nowadays. To construct EA aggregates before 1999, we have used the same aggregation procedure that is currently used by the ECB, namely, GDP-weighted averages (measured in units of PPP at 1995 prices) of: (i) the relevant short-term intervention interest rates in the member countries of the EA, (ii) real output growth (measured by monthly growth rates of the *Industrial Production Index*, IPI), and (iii) inflation rates (measured by the *Harmonised Price Index*, HCPI).

Accordingly, our econometric approach is a two-stage one. In the first stage, a Taylor rule is estimated for the EA, as if a surrogate ECB was exerting monetary policy control during the whole period under study. Thus, the residuals of this reaction function will be interpreted as monetary-policy (interest rate) shocks. A novel feature of this exercise is that, instead of estimating the conventional (forward-looking) linear Taylor rule popularized by Clarida *et al.* (1998, 2000), we estimate a nonlinear rule which accounts for nonlinearity in the Phillips curve, given the ample evidence in favour of this hypothesis.<sup>3</sup> In the second stage, the shocks from the Taylor rule are then used as explanatory variables in a MMS model in order to test the existence of asymmetric effects of *unanticipated* changes in the monetary-policy stance on real output growth depending on the phase of the business cycle.

Proceeding in this fashion, we obtain two interesting results. First, evidence is found in favour of *state* asymmetries at the aggregate level in the

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<sup>3</sup>For instance, convexity in the short-run inflation-output gap trade-off arises under the traditional Keynesian assumption that nominal wages are flexible upwards and rigid downwards, giving rise to a convex *aggregate supply* schedule (see, e.g., Baily, 1978). More recently, Akerlof *et al.* (1996) have further elaborated on that argument claiming that even a long-run trade off exists at very low rates of inflation due to the existence of money illusion on the part of the workers when there is price stability. Laxton *et al.* (1995, 1999) and Gerlach (2000) have presented evidence supporting a convex Phillips curve in the inflation-output gap space for several European countries and the US. Schaling (1999), Orphanides and Wieland (2000) and Dolado *et al.* (2005) are the first papers to consider this type of nonlinearity in the derivation of optimal monetary policy rules.

EA, whereby interest rate shocks have larger effects in recessions than in expansions. Second, we find that interest rate shocks also affect the transition probabilities from one cyclical phase to another in a different way, namely, an interest rate rise in a boom is less effective in allowing the economy to become less expansionary than an equally-sized interest rate cut in helping the economy to escape from a slump.

The remainder of the chapter is organised as follows. Section 2 provides a brief overview of the literature dealing with asymmetries in the effects of monetary-policy shocks on the real side of the economy. In Section 3, a nonlinear forward-looking Taylor rule is estimated for the EA in order to derive these shocks. Section 4 offers a brief explanation of the basics of the MMS methodology which is used throughout the rest of the chapter, and presents results for the effects of the shocks on real output growth rate in a model with constant transition probabilities. Section 5 relaxes the previous assumption by allowing the transition probabilities to be directly affected by the shocks. Finally, Section 6 draws some conclusions.

## 2 Related literature

There are three main types of asymmetries which have been discussed in the literature about the effects of *unanticipated* monetary policy changes on real output: (i) the traditional *Keynesian* asymmetry, associated with the *sign* of the monetary shocks, (ii) the standard *menu cost* asymmetry related to the *size* of those shocks, and (iii) the *state* asymmetry whereby the effects of monetary shocks on output depend on the phases of the business cycle.

One can find a wide range of theoretical contributions in the literature that provide microfoundations for these asymmetries. In relation to the *sign* asymmetry, its rationale relies upon the nominal stickiness properties of menu costs and has been examined, among many others, by Akerlof and Yellen (1985). As for the *size* asymmetry, Ball and Romer (1989, 1990), Caballero and Engel (1992) and Tsiddon (1991), *inter alia*, have analysed S-s threshold-type price adjustment rules which lead to convex aggregate supply curves, as in the standard Keynesian framework.

More recently, the possibility of having a *hybrid* asymmetry, according to which only *small negative* shocks affect real output, has been considered

as well in models which combine dynamic menu-costs with a positive trend inflation rate. As Ball and Mankiw(1994) have argued, the underlying explanation for that type of asymmetry is that, in the face of a positive trend inflation rate, *small negative* shocks should bring the actual price closer to the optimal value and the opposite should be expected when shocks are positive, either large or small. Consequently, in this case, firms will not adjust their prices and, therefore, real effects will take place.

Empirical support for both types of asymmetries is well documented in the literature. On the one hand, Cover (1992), De Long and Laurence (1998) and Karras (1996) have found favourable evidence for the *Keynesian* asymmetry in the US and a number of European countries. On the other hand, Ravn and Sola (1996) and María-Dolores (2001) find strong evidence for both the *Keynesian* and *menu-cost* asymmetries in the US and Spain, respectively.

Our aim in this chapter is restricted to the empirical analysis of *state* asymmetries, namely, whether unanticipated changes in the monetary policy stance affect real output differently in upturns than in downturns. Although this type of asymmetry has received far less attention in the literature than the other two asymmetries, there are at least two compelling arguments which make them relevant. First, the previously discussed price- adjustment models leading to a convex aggregate supply schedule could be re-interpreted as implying that monetary policy will have stronger real effects during recessions, when output is below its long-run level, than in expansions, when the aggregate supply curve is almost vertical. And secondly, there is a broad class of models which provide support for this type of asymmetry by explicitly modelling the credit or lending channel of the monetary transmission mechanism. According to this interpretation, if financial markets face information asymmetries, credit and liquidity may be readily available in booms whilst agents may find it harder to obtain funds in slumps. Therefore, it is likely that monetary policy will have stronger effects on the consumption and investment decisions during upturns than during downturns. This is the mechanism derived from the extensive research on financial market imperfections, including agency costs and debt overhang models, developed, *inter alia*, by Bernanke and Gertler (1989), Gertler (1988), Gertler and Gilchrist (1994), Kiyotaki and Moore (1998) and Lamont (1993). As for the empirical support of this type of asymmetry using the MMS modelling approach, to our knowledge, the only available studies are those by Ravn and Sola (1996) and

García and Schaller(1995) who provide favourable evidence for *state* asymmetries in the US , and Dolado and María-Dolores(2001) who find them in Spain.<sup>4</sup> Our contribution here will rely heavily upon the methodological approach proposed in these papers.

### 3 Estimation of a monetary policy reaction function

In this section, following the arguments in Dolado *et al.* (2005), we estimate a nonlinear *forward-looking* Taylor rule for the ECB. The setup is as follows. Let us suppose that the policymaker sets the nominal interest rate,  $i$ , with the goal of minimizing inflation deviations from a target,  $\mathfrak{e}=\pi-\pi^*$ , and the output gap,  $\tilde{y}$  , in every period. Assuming a quadratic per-period loss function in inflation and output performance,  $L(\mathfrak{e}_t, \tilde{y}_t) = \frac{1}{2}[\mathfrak{e}_t^2 + \lambda\tilde{y}_t^2]$ , and a fixed discount rate  $\delta$ , the policymaker's objective in period  $t$  is to minimise the expected present discounted value of the per-period losses:

$$E_t \sum_{s=0}^{\infty} \frac{1}{2} \delta^s L(\mathfrak{e}_{t+s}, \tilde{y}_{t+s}), \quad (1)$$

subject to the following two equations describing the evolution of the economy:

$$\pi_{t+1} = \pi_t + \alpha f(\tilde{y}_t) + u_{\pi,t+1}, \quad (2)$$

with

$$f(\tilde{y}_t) = \tilde{y}_t + \phi\tilde{y}_t^2, \quad \tilde{y}_t > \frac{-1}{2\phi}, \quad (3)$$

and

$$\tilde{y}_{t+1} = \beta\tilde{y}_t - \zeta r_t + u_{y,t+1}, \quad (4)$$

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<sup>4</sup>There is, however, a large literature on asymmetries in business cycles considered from a univariate perspective. See, e.g., Neftci (1984), Beaudry and Koop (1993), Huh (1993) and McQueen and Thorley (1993).

where  $E_t$  is the conditional expectations operator,  $\delta$  and  $\beta \in [0, 1)$ , and  $u_{\pi,t+1}$  and  $u_{y,t+1}$  are zero-mean normally distributed shocks.

Equation (2) represents an accelerationist Phillips curve, or aggregate supply (*AS*) schedule, where where the output gap enters in a nonlinear way, as defined in equation (3). Note that the conventional linear *AS* schedule is recovered when  $\phi = 0$ , and that the function is convex (concave) if  $\phi > 0$  ( $< 0$ ). As any *AS* schedule, it is assumed to be increasing ( $1 + 2\phi\tilde{y} > 0$ ) for realistic values of  $\phi$  and  $\tilde{y}_t$ .<sup>5</sup> Equation (4), in turn, represents an *IS* schedule where the output gap exhibits sluggish adjustment, and depends on the real interest rate ( $r_t = i_t - E_t\pi_{t+1}$ ). Notice that the real interest rate affects output with one-period lag and, therefore, affects inflation with a two-period lag. This timing convention, borrowed from Svensson (1997), is in line with the extensive literature on the transmission mechanism of monetary policy which establishes that an innovation in monetary policy leads to a change in output in the short run, with inflation only changing slowly later on (see e.g., Christiano *et al.*, 1999).

Totally differentiating (1) with respect to  $i_t$ , subject to (2) – (4) , yields the following Euler equation:

$$\lambda E_t \tilde{y}_{t+1} + \lambda \delta \beta E_t \tilde{y}_{t+2} + \delta \alpha E_t \pi_{t+2} (1 + 2\phi \tilde{y}_{t+1}) = 0. \quad (5)$$

Using (4) to replace  $E_t \tilde{y}_{t+2}$  in terms of  $E_t \tilde{y}_{t+1}$  and  $E_t r_{t+1}$ , and solving for the optimal value of  $i$  (denoted as  $i^*$ ) in period  $t$ , implies that the policymaker should set  $i_t^*$  according to the following reaction function:

$$i_t^* = c_1 E_{t-1} \pi_{t+1} + c_2 E_{t-1} \tilde{y}_t + c_3 E_{t-1} (\pi_{t+1} \tilde{y}_t), \quad (6)$$

where the  $c_i$ 's coefficients are functions from the set of structural parameters ( $\delta, \alpha, \lambda, \phi, \zeta$  and  $\beta$ ).<sup>6</sup>

Our new monetary policy rule in (6) looks like a standard linear Taylor rule except for the last term, namely, the expected interaction of current output and future inflation. The presence of the *interaction* term in the Euler equation above is quite intuitive. Take, for example, the case where inflation is expected to be above its target at period  $t + 1$  by one percentage

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<sup>5</sup>This is the case for the range of values of  $\tilde{y}$  and the estimated value of  $\phi$  in our sample.

<sup>6</sup>It can be shown that  $c_1 = 1 + \alpha/\lambda\zeta\beta$ ,  $c_2 = (1 + \delta\beta^2)/\delta\zeta\beta$ ,  $c_3 = 2\phi\alpha/\lambda\zeta\beta$ .

point. Then, the real interest rate will be below its equilibrium value at period  $t$  which, in turn, causes a higher output gap at  $t + 1$  and higher inflationary pressure at  $t + 2$ . In the linear case, the policymaker increases the interest rate by  $c_1$ . However, if the Phillips curve is convex ( $\phi > 0$ ), then the future inflationary pressure caused by the higher output gap will turn out to be larger than in the linear case. The policymaker, anticipating this higher pressure, captured by the *interaction* term, will react more forcefully by implementing a larger rise in the interest rate, since in this case  $c_3 > 0$ .

To estimate the policy rule, as is customary, we replace the expectations in (6) by their realized values, yielding :

$$i_t^* = cns + c_1 \mathbf{e}_{t+1} + c_2 \tilde{y}_t + c_3 (\mathbf{e}_{t+1} \tilde{y}_t) + \xi_t. \quad (7)$$

From the viewpoint of testing asymmetries in the monetary policy reaction function what really matters is the  $c_3$  coefficient. This is the only coefficient which embodies information on the nonlinear Phillips curve, so that the restriction  $\phi = 0$  implies  $c_3 = 0$ . Indeed, it is straightforward to check that the ratio  $c_3/2(c_1 - 1)$  yields a direct estimate of  $\phi$ . Hence testing  $H_0: \phi = 0$  is equivalent to testing  $H_0 : c_3 = 0$  as long as  $c_1$  is different from unity.<sup>7</sup> Since (7) is linear in the coefficients, the key advantage of testing directly  $H_0 : c_3 = 0$  is that it does not require estimating a nonlinear model in the parameters. As for the error term in (7), it is defined as:

$$\begin{aligned} \xi_t = & -[c_1(\mathbf{e}_{t+1} - E_{t-1}\mathbf{e}_{t+1}) + c_2(\tilde{y}_t - E_{t-1}\tilde{y}_t) + \\ & + c_3(\mathbf{e}_{t+1}\tilde{y}_t - E_{t-1}(\mathbf{e}_{t+1}\tilde{y}_t))], \end{aligned} \quad (8)$$

where the term in brackets is a linear combination of forecast errors and therefore orthogonal to any variable in the information set available at  $(t-1)$ . Note that our specification is very similar to the one popularised by Clarida *et al.* (1998, 2000) except for the inclusion of the interaction term between inflation and the output gap. For estimation purposes, however, we will use a slight modifications of equation (7) concomitant to the use of data with a monthly frequency. First, in accord with most of the empirical literature, we take one year ( $k = 12$ ) to be the horizon used by the central bank in

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<sup>7</sup>This is the case in our estimation below.



forecasting inflation. And, secondly, as is also conventional, we use a lagged dependent variable to capture interest-rate smoothing for which there are several motivations in the literature. For this reason, the estimated rule will be the following partial-adjustment model:

$$i_t = \rho_1 i_{t-1} + (1 - \rho_1) i_t^* + \xi_t. \quad (9)$$

As is conventional in the estimation of Taylor rules, the estimation method relies upon the choice of a set of instruments,  $Z_t$ , from the set of variables within the central bank's information set, such as lagged variables that help forecast inflation and output or any other contemporaneous variables that are uncorrelated with the policy rule shock,  $\xi_t$ .<sup>8</sup> Then, the Generalized Method of Moments (*GMM*) can be used to estimate the parameter vector in (7) by exploiting the set of orthogonality conditions  $E(\xi_t/Z_t) = 0$ . Since the composite disturbance  $\xi_t$  has an MA( $k$ ) representation, due to the overlapping nature of the forecast errors, the Newey-West weighting var-cov matrix is used to implement GMM. Finally, Hansen's (1982)  $J$  test is used to test the overidentification restrictions.

As explained in the Introduction, equation (9) is estimated using monthly data for the Euro-area.<sup>9</sup> The sample period, 1996(01)-2003(12), has been chosen on the basis of selecting a sufficiently large sample size (96 observations) and a homogeneous recent spell where the implementation of monetary policy by the national central banks before 1999 was fairly similar to that conducted by the ECB later on. The short-term intervention interest rate is chosen to be a weighted average of short-term intervention interest rates for the Euro-area before 1999 and the Euro-area interest rate after January 1999. Inflation is measured through the HCPI inflation rate and output through (logged) IPI since this is the only available measure of real output on a monthly basis. All variables are seasonally adjusted and the IPI has been corrected from calendar effects. To obtain a measure of output gap, we detrend the (log of) IPI using the HP filter with a coefficient of 14.800.<sup>10</sup>

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<sup>8</sup>The list of instruments is: a constant term, two lags of the interest rate, six lags of the inflation rate, six lags of the output gap, four lags of the interaction between inflation and output gaps and two lags of a (logged) raw materials price index.

<sup>9</sup>The data source is the OECD data base ([www.sourceOECD.org](http://www.sourceOECD.org))

<sup>10</sup>This coefficient is the default value in E-Views 5.0 for monthly data. Other values in that range led to similar results, as did the residuals from adjusting a cubic trend to logged output or from applying the band-pass filter of Baxter and King (1999).

Finally, as regards the inflation target,  $\pi^*$ , we use the official ECB target of 2%.

To get some preliminary evidence on the key channel for the nonlinear Taylor rule in (7), Table 1 reports the results from the OLS estimation of the nonlinear Phillips curve proposed in (3).<sup>11</sup> The change in inflation at time  $t$ ,  $\Delta\pi_t$ , has been regressed on  $f(\tilde{y}_{t-1})$  to estimate the parameters  $\alpha$  and  $\phi$ . A positive and statistically significant estimate of  $\phi$  implies a convex Phillips curve. As can be observed, there is favourable evidence to such a hypothesis. To stress it, Figure 1 depicts the scatter plot of lagged output gap (horizontal axis) against the change in inflation (vertical axis), together with the fitted quadratic function, where it becomes clear than the curve is convex.

[TABLE 1]

[FIGURE 1]

Table 2, in turn, displays the estimated coefficients of (9). The coefficient on the lagged interest rate ( $\rho_1$ ) is estimated to be 0.83, indicating a fairly sluggish adjustment, in line with the available estimates in the literature. The point estimate of  $c_1$  is always above unity which, as argued by Clarida et al. (1998, 2000), implies an inflation-stabilising policy rule. Moreover, the estimate of  $c_2$  is also strongly significant pointing out to a response of the ECB to real activity, since the current output gap is a leading indicator of future inflationary pressures. The most relevant result, however, is the positive sign and high statistical significance of the estimate of  $c_3$ , yielding strong support to the existence of a nonlinear (convex) Phillips curve in the EA. This result is in line with a similar finding by Dolado *et al.* (2005) for the European Union using quarterly data for the period 1984-2001. Notice that an estimate of  $\phi$  can be retrieved from the ratio  $c_3/2(c_1 - 1)$  which yields 0.51, i.e., a value fairly similar to that reported in Table 1.<sup>12</sup> Finally, it is worth noticing that the  $p$ -value of the  $J$ -test (denoted as  $p$ - $J$  in Table 2) does not reject the over-identifying restrictions.

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<sup>11</sup>In this and the remaining Tables, the superscripts \*\*, \*, and † denote significance at the 1%, 5% and 10% levels, respectively.

<sup>12</sup>Indeed, using the delta method to compute 95% intervals of  $\phi$ , we cannot reject the null that the difference between both values is zero.

Figure 2 depicts the short-term interest rate (solid line) together with the Taylor rule predictions (dotted line). Note that the performance of the non-linear Taylor rule is particularly striking since 1999, when the ECB started its operations. However, before 1999, the rule slightly underpredicts although it captures the fall in the interest rate.

[TABLE 2]

[FIGURE 2]

## 4 Markov switching models for real output growth

In this section, we introduce the basic principles of the econometric modelling approach which is applied to analyse the existence of asymmetries in the effects of unanticipated monetary policy changes on real activity. We start by presenting a brief overview of the basics of the MMS methodology in relation to the conventional univariate approach popularized by Hamilton's (1989). Next, we explain how the *state* asymmetric effects of monetary policy can be tested in this framework.

### 4.1 Extended Markov- Switching model including interest-rate shocks

In line with the approach advocated by Ravn and Sola (1996), Garcia and Schaller (1995) and Dolado and Maria-Dolores (2001) to investigate the presence of state asymmetries in the effects of monetary policy shocks on output, we rely upon a MMS model. In this model, real output growth is allowed to be affected by the interest rate shocks,  $\xi_t$ , obtained from equation (9) so that their effects depend on the state of the economy.

As is well known, Hamilton's MS univariate approach relies upon the assumption that the actual state of the business cycle, i.e., a recession ( $r$ ) or an expansion ( $e$ ), is determined by an unobserved latent random variable which follows a Markov process. In Hamilton's (1989), the growth rate of output ( $\mu$ ) in the US economy was assumed to follow an  $AR(p)$  process whose

unconditional mean ( $\mu$ ) and autoregressive coefficients ( $\phi_i$ 's) were allowed to vary as function of whether the economy was in an expansion ( $\mu_e, \phi_{ie}$ ) or in a recession ( $\mu_r, \phi_{ir}$ ). We follow the same approach here. Yet, to estimate the asymmetric effects of monetary policy shock ( $\xi$ ) on output growth, the  $AR(p)$  model is generalized to allow for the varying effects of the shocks in the following way:

$$\begin{aligned} \Delta y_t = & \phi_1 \Delta y_{t-1} + \dots + \phi_p \Delta y_{t-p} + \mu_r (1 - \phi_1 - \dots - \phi_p) + \\ & \Delta \mu (S_t - \phi_1 S_{t-1} - \dots - \phi_p S_{t-p}) + \\ & \beta_{1r} \xi_{t-1} + \Delta \beta_1 S_{t-1} \xi_{t-1} + \dots + \beta_{pr} \xi_{t-p} + \Delta \beta_p S_{t-p} \xi_{t-p} + \sigma \eta_t \end{aligned} \quad (10)$$

where  $\Delta y$  is the (monthly) IPI growth rate,  $\Delta \mu = \mu_e - \mu_r$ ,  $S_t$  is the state variable and  $\eta_t$  is distributed  $N(0, 1)$  normalized by its standard deviation,  $\sigma$ . Further,  $\Delta \beta = \beta_e - \beta_r$ , and  $\beta_r$  and  $\beta_e$  are the coefficients on the shocks in recessions and expansions, respectively. Therefore, the chosen specification implies that the effects of the shocks on output growth depend on the business cycle phase that the economy was undergoing at the time the shock took place.<sup>13</sup> In order to account both for the lagged effects of shocks on output growth and to define the shocks as predetermined variables in (10) (see Filardo 1994), we set the contemporaneous effect of  $\xi_t$  on  $\Delta y_t$  equal to zero. Finally, note that the autoregressive coefficients,  $\phi_i$ , in (10) have been assumed to be independent of  $S_t$ , since a LR test (with a p-value of 0.27) did not reject the null hypothesis of parameter constancy in this subset of coefficients against state dependence.

As is conventional, the state variable in the model,  $S_t$ , is assumed to follow a discrete-time Markov process which is characterized by the following transition probability matrix  $\Pi$ :

$$\begin{pmatrix} p_{rr} & p_{er} \\ p_{re} & p_{ee} \end{pmatrix} = \begin{pmatrix} p_{rr} & 1 - p_{ee} \\ 1 - p_{rr} & p_{ee} \end{pmatrix}, \quad (11)$$

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<sup>13</sup>If instead of specifying the effects of the shocks as  $\Delta \beta_i S_{t-i} \xi_{t-i}$  ( $i = 0, \dots, p$ ) in (10), we were to introduce them in the form  $\Delta \beta_i S_t \xi_{t-i}$ , then the effects would depend on the *current* state of the economy rather on the state at the time the shock took place. Estimation of this alternative specification of the state dependence yields fairly similar results to those presented in Table 3 below and therefore are omitted.

where:

$$p_{ij} = \Pr(S_t = j/S_{t-1} = i), \text{ with } \prod_{j=r}^e p_{ij} = 1 \text{ for all } i, \quad (12)$$

such that  $p_{ij}$  is the probability of going from state  $i$  to state  $j$  (e.g.,  $p_{re}$  is the probability of going from a recession to an expansion, etc.). Initially, we assume that the transition probabilities are constant over time and are determined by the following simple logistic specifications:

$$p_{rr} = \Pr(S_t = r/S_{t-1} = r) = \frac{\exp(\theta_r)}{1 + \exp(\theta_r)} \quad (13)$$

$$p_{ee} = \Pr(S_t = e/S_{t-1} = e) = \frac{\exp(\theta_e)}{1 + \exp(\theta_e)}, \quad (14)$$

where  $\theta_r$  and  $\theta_e$  denote the parameters that determine the probabilities of being in a recession and in an expansion, respectively.

As Hamilton (1989) showed, the above assumptions allow us to obtain a sequence of joint conditional probabilities  $\Pr(S_t = i, \dots, S_{t-s} = j / \Phi_t)$ , which are the probabilities that the output growth series is in state  $i$  or  $j$  ( $i, j = r, e$ ) at times  $t \dots t - s$ , respectively, conditioned by the information available at time  $t$ . By adding those joint probabilities we can obtain the so-called (*smoothed*) *filter* probabilities, namely, the probabilities of being in states  $r$  or  $e$  at time  $t$ , given information available at time  $t$ :

$$\Pr(S_t = j / \Phi_t) = \prod_{i=r}^e \dots \prod_{j=r}^e \Pr(S_t = i, \dots, S_{t-s} = j / \Phi_t), i, j = e, r, \quad (15)$$

where  $\Phi_t$  is the set of available information in period  $t$ . The smoothed filter probabilities provide information about the regime in which the series are most likely to have been in time  $t$  at every point in the sample. Therefore, they turn out to be very useful tools for dating phase switches.

Table 3 report the estimates of the coefficients in (10), where four lags ( $p = 4$ ) have been chosen on the basis of the Akaike (AIC) and Bayesian (BIC) information criteria. The first regime corresponds to a contractionary phase with a monthly growth rate of  $-0.29\%$  ( $-3.36\%$  annually). while the second

regime clearly corresponds to an expansionary phase with a monthly growth rate of 0.38% (4.56% annually). As regards the probabilities of remaining in each regime, they are estimated to be 0.89 for a recession and 0.94 for an expansion<sup>14</sup>. These probabilities imply mean durations of 9.1 (slumps) and 16.7 (booms) pointing out that that recessions over the sample period tended to be shorter than expansions<sup>15</sup>. Figure 3, in turn, depicts the smoothed probabilities of a recession together with the monthly IPI growth rate. As can be observed, these probabilities tend to have a strong inverse correlation with the evolution of IPI growth so that they are high (low) when the growth rate is negative (positive)<sup>16</sup>. Finally, from the viewpoint of this chapter, the most relevant finding is that the sizes of the (negative) slopes of the shocks seem to be larger in recessions than in expansions, pointing out to the existence of state asymmetries. Indeed, a LR test of  $H_0 : \beta_{ir} = \beta_{ie} (i = 1, 4)$  rejects the null hypothesis of symmetric effects with a p-value of 0.008. This feature is illustrated by Figure 4 which depicts the impulse-response functions of output growth to a one-standard-deviation shock in  $\xi_t$ . It becomes clear the output effects of an unanticipated rise in the interest rate are much stronger in a contractionary than in an expansionary phase or, alternatively, that an unanticipated rise of the interest rate in a boom is bound to reduce real output growth by less than an equally-sized reduction of the interest rate would increase output in a slump.

[TABLE 3]  
[FIGURE 3]

[FIGURE 4]

Finally, the estimated MMS model can be used to forecast (monthly) IPI growth by writing (10) in the vector AR(4) companion form:

$$\begin{aligned} \Delta \mathbf{y}_t = & \phi(L)\Delta \mathbf{y}_{t-1} + \boldsymbol{\mu}(1 - \phi(L)) + \Delta \boldsymbol{\mu} S_t(1 - \phi(L)) \\ & + \boldsymbol{\beta}(L)\boldsymbol{\xi}_{t-1} + \Delta \boldsymbol{\beta}(L) S_{t-1}\boldsymbol{\xi}_{t-1} + \boldsymbol{\sigma}\boldsymbol{\eta}_t \end{aligned} \quad (16)$$

---

<sup>14</sup>The probabilities have been computed as in (13) and (14).

<sup>15</sup>In Table 3, mean durations of recessions and expansions are denoted by  $d_r$  and  $d_e$ , respectively.

<sup>16</sup>The correlation coefficient is  $-0.80$

such that,

$$\Delta y_{t+h/t} = e_1' [\phi(L)\Delta y_{t+h-1/t} + \mu(1 - \phi(L)) + \Delta\mu S_{t+h/t}(1 - \phi(L)) + \beta(L)\xi_{t+h-1/t} + \Delta\beta(L)S_{t+h-1/t}\xi_{t+h/t}] \quad (17)$$

where  $L$  is the lag operator,  $\phi(L)$  and  $\beta(L)$  are the corresponding matrices in the companion form and  $e_1' = (1, 0, \dots, 0)$ . The autoregressive forecasts of  $S_{t+h/t} = \Pi^h S_t$  and  $\xi_{t+h/t}$  are obtained from the dynamic forecasts of the Taylor rule in (9) using ARIMA processes to compute estimates of  $\hat{\epsilon}_{t+h+k}$ ,  $\hat{\eta}_{t+h}$  and their interactions ( $\hat{\epsilon}_{t+h+k} \hat{\eta}_{t+h}$ ). Table 4 presents the forecasts and forecast errors of the (monthly) growth rate of IPI for the period 2003:01 to 2003:12. The *RMSE* of the forecasts is 0.114 which fares very well in comparison with the *RMSE* obtained from simple AR (4) (0.512) and AR(12) models (0.457), and also in relation to a re-estimated univariate MS model, like the one in Table 3, but excluding the effects of the shocks (0.266).<sup>17</sup>

[TABLE 4]

## 5 Effects of Monetary Policy on State Switches

Whereas in the previous section we allowed for state dependence in the effects of interest- rate shocks on output growth, the transition probabilities from one phase to another were assumed to be independent of those shocks. Thus, while we were able to test whether shocks had different incremental effects on output in each state, we were not able to examine whether those shocks might have a further effect on output growth by directly affecting the probability of a state switch. In this section, we address this issue by allowing those probabilities to depend directly on the shocks.<sup>18</sup> Hence, the logit functions in (13) and (14) are replaced by:

$$p_{rr} = \Pr(S_t = r / S_{t-1} = r) = \frac{\exp(\theta_{or} + \theta_{1r}\xi_{t-1} + \theta_{2r}\xi_{t-2})}{1 + \exp(\theta_{or} + \theta_{1r}\xi_{t-1} + \theta_{2r}\xi_{t-2})} \quad (18)$$

<sup>17</sup>Hansen's (1992) statistic to test constant linear autoregressive coefficients under the null against the last MS model, excluding the shocks, yields a p-value of 0.017, therefore, rejecting the null.

<sup>18</sup>The maximization algorithm with variable transition probabilities is considered in Filardo (1994).

$$p_{ee} = \Pr(S_t = e/S_{t-1} = e) = \frac{\exp(\theta_{oe} + \theta_{1e}\xi_{t-1} + \theta_{2e}\xi_{t-2})}{1 + \exp(\theta_{oe} + \theta_{1e}\xi_{t-1} + \theta_{2e}\xi_{t-2})}, \quad (19)$$

where only two lags of  $\xi_t$  has been chosen in (18) and (19) to keep the number of parameters manageable. Further, in order to isolate the effect of the shocks from the linear effects examined above, the coefficients on the latter terms ( $\beta_{ir}$  and  $\beta_{er}$ ) are constrained to be zero, as in García and Schaller (1995) and Dolado and María-Dolores (2001).<sup>19</sup> Notice that since the probability of remaining in a recession (expansion) is increasing in the  $\theta_{ir}(\theta_{ie})$  parameters, we should expect  $\theta_{1r}$  and  $\theta_{2r}$  to be positive, and  $\theta_{1e}$  and  $\theta_{2e}$  to be negative, when considering an interest rate rise. In other words, an increase in the interest rate should reduce the probability of remaining in an expansion and increase the probability of remaining in a recession.

The estimates of the coefficients in the MMS model with variable transition probabilities are reported in Table 5, where it can be observed that the signs of the  $\theta_{ir}$  and  $\theta_{ie}$  coefficients are in agreement with the above interpretation. Furthermore, the restriction stemming from ignoring the linear effects of shocks has a limited effect on the estimates of the probabilities  $p_{ee}$  and  $p_{rr}$ , since the estimated intercepts ( $\theta_{0r}, \theta_{0e}$ ) yield, according to (13) and (14),  $p_{ee} = 0.90$  and  $p_{rr} = 0.87$ , namely, close values to the respective probabilities (0.94 and 0.89) reported in Table 3.

[TABLE 5]

To ascertain the effects of interest rate shocks on the transition probabilities, we use the following experiment. Suppose that the ECB were to implement a negative (expansionary) interest rate shock of  $x$  basis points in two consecutive months (from  $t$  to  $t + 2$ ), in agreement with the number of lags with which  $\xi_t$  appears to affect  $p_{rr}$  and  $p_{ee}$  in (18) and (19). Then, the question would be: *How would such a reduction in the interest rate affect*

---

<sup>19</sup>Furthermore, attempts to estimate an encompassing model with state dependent probabilities and state vaying effects of the shocks, as in (10), failed to achieve convergence of the Filardo's algorithm.



*the transition probability from a recession to an expansion?* Likewise, if, instead, a positive (contractionary) interest rate shock of identical magnitude were to be considered: *How would such a rise in the interest rate affect the probability of a converse switch?*

As an illustration of the proposed simulation, Table 6 shows the changes in  $p_{er}$  ( $p_{re}$ ) in response to a positive(negative) interest rate shock of 100 b.p. ( $-100$  b.p.). It is found that an unanticipated interest-rate cut of such a magnitude will increase the probability of getting out from a recession ( $p_{re}$ ) from 0.13 to 0.24 whereas an unanticipated rise in the interest rate will increase the probability of entering a recession ( $p_{er}$ ) from 0.10 to 0.16. Note that, in accord with the stronger real effects of monetary policy during recessions found before, the increase in the probability of escaping a recession in response to a cut in interest rates is about twice as large as that the reduction in the probability of entering a recession in response to a rise in interest rates.

[TABLE 6]

## 6 Conclusions

In this chapter, we have investigated the possibility of asymmetric effects of monetary policy shocks on real output growth in the EA depending on the business cycle phase of the economy. This type of asymmetric effects are known as *state* asymmetries, according to which the effects of monetary policy shocks on real activity may be stronger in recessions than in expansions. The rationale for these asymmetries stems from an extensive theoretical research which stresses financial market imperfections, including models which deal with credit crunches and debt overhang, in the transmission mechanism from monetary policy on output.

Our test of this asymmetry in the EA, using monthly data over the period 1996-2003, relies on a Markov-switching model of real output (IPI) growth augmented with predetermined variables whose coefficients or the transition probabilities are allowed to depend on the latent state variable which identifies the cyclical phases. Our results here offer some support for the previous

hypothesis. In particular, we find that monetary policy shocks, measured as shocks to the short-term interest rate obtained from a forward-looking Taylor rule, have significantly larger effects during recessions than during expansions, and that unanticipated interest- rate cuts help about twice more to escape a recession than a corresponding rise in interest rates would help to cool down the economy when it is undergoing an expansion.

Finally, as a by-product of our analysis, we have found significant evidence of nonlinearity in the policy rule of the ECB, in the sense that it has tended to intervene more vigorously when inflation and output move together above their targets than what a linear Taylor rule would predict, particularly after 1999. This result could be interpreted in terms of the existence of a convex Phillips curve in the EA.

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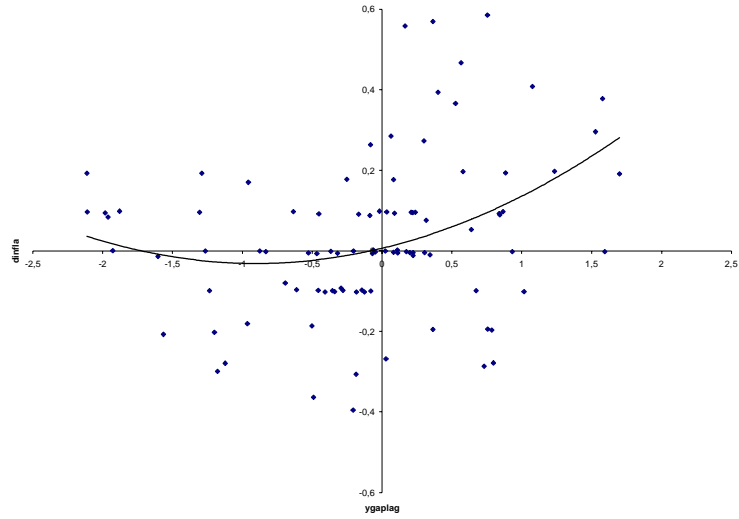


Figure 1: Phillips Curve for the Euro-Area

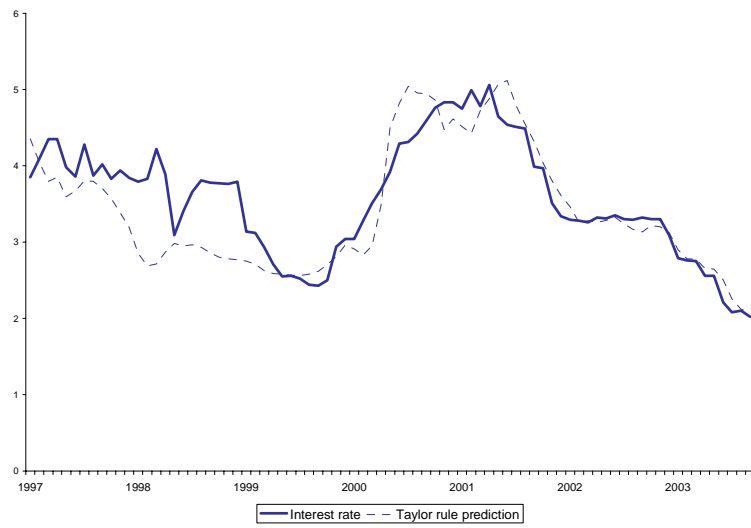


Figure 2: Interest rate for the Euro-Area and Taylor rule predictions (monthly data)



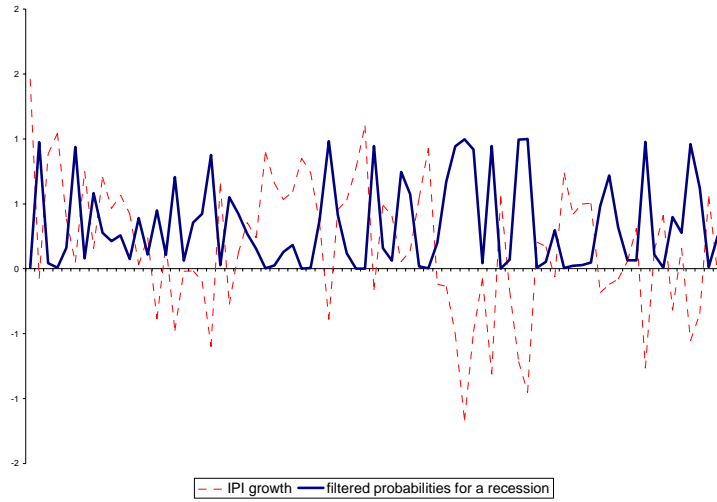


Figure 3: Monthly IPI growth rate and smoothed probabilities for a recession

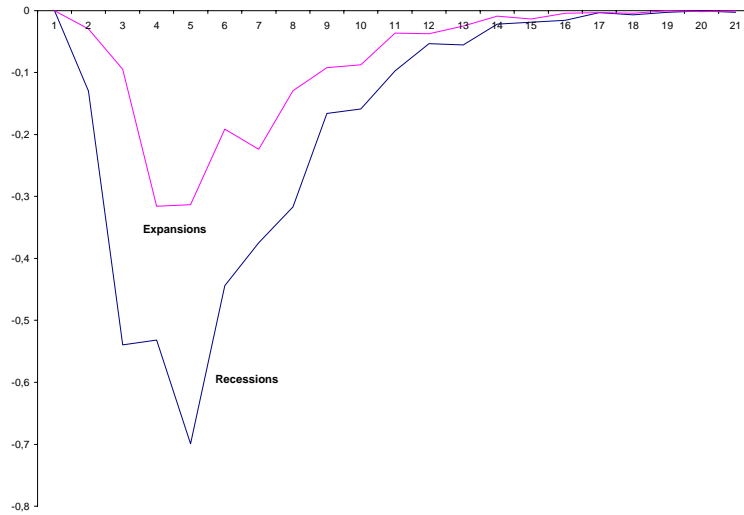


Figure 4: Impulse-Response function to an unanticipated increase in interest rate (monthly data)

**Table 1: Estimated Phillips Curve for Euro-zone**

Coefficient	Estimate
$\alpha$	0.084* (0.024)
$\phi$	0.583* (0.122)

*Note:* The figures in parenthesis are White's standard errors.

**Table 2. Estimated Taylor rule for Euro-zone**

Coefficient	Estimate
<i>cnst</i>	0.54** (0.10)
$\hat{\rho}_1$	0.83** (0.03)
$\mathbf{b}_1$	1.40** (0.46)
$\mathbf{b}_2$	0.11** (0.02)
$\mathbf{b}_3$	0.41** (0.06)
$\mathbf{b}_\varepsilon$	0.25
$p - J$	0.086

*Notes:* The figures in parentheses are standard errors.  
p-J is the p-value of the J-test over-identifying restrictions

**Table 3**  
**MMS model for IPI growth**  
**Dependent variable  $\Delta y_t$**

Coefficient	Estimate
$\mu_r$	-.29* (.15)
$\mu_e$	.38** (.12)
$\phi_1$	.15* (.07)
$\phi_2$	.39† (.21)
$\phi_3$	.30† (.17)
$\phi_4$	-.23 (.11)
$\beta_{1r}$	-.13 (.11)
$\beta_{1e}$	-.03 (.07)
$\beta_{2r}$	-.52† (.27)
$\beta_{2e}$	-.09 (.21)
$\beta_{3r}$	-.40* (.19)
$\beta_{3e}$	-.29** (.13)
$\beta_{4r}$	-.37* (.17)
$\beta_{4e}$	-.22** (.08)
$\sigma$	.102** (.03)
$\theta_r$	2.09** (.045)
$\theta_e$	2.75** (.67)
$d_r$	9.09
$d_e$	16.67
Log-Likelihood	75.67

Note: standard errors in parentheses

<b>Table 4</b>						
<b>Forecasts of the IPI (monthly) growth rate (%): 2003</b>						
Month	<b>03(01)</b>	<b>03(02)</b>	<b>03(03)</b>	<b>03(04)</b>	<b>03(05)</b>	<b>03(06)</b>
Current	-0.0790	0.0587	0.3142	-0.7665	0.1407	0.4138
Forecast	-0.0917	0.0602	0.2530	-0.5551	0.2550	0.2635
For. Error	0.0127	-0.0015	0.0612	0.2114	-0.1080	0.1502
Month	<b>03(07)</b>	<b>03(08)</b>	<b>03(09)</b>	<b>03(10)</b>	<b>03(11)</b>	<b>03(12)</b>
Current	-0.3204	0.1585	-0.5547	-0.3477	0.5589	-0.047
Forecast	-0.1607	0.2902	-0.6417	-0.2592	0.4569	-0.122
For. Error	-0.1597	-0.1317	0.0870	-0.0884	0.1010	0.0758

Table 5  
 Markov Switching Model with variable transition  
 probabilities

Coefficient		Estimate
$\mu_r$		-.26** (.05)
$\mu_e$		.35** (.01)
$\phi_1$		.37* (.14)
$\phi_2$		.23* (.12)
$\sigma$		.13** (.003)
$\theta_{0r}$		1.87** (.32)
$\theta_{1r}$		.47** (.05)
$\theta_{2r}$		.25** (.05)
$\theta_{0e}$		2.18** (.73)
$\theta_{1e}$		-.34** (.08)
$\theta_{2e}$	-	-.18** (.05)
Log-Likelihood		77.21

Note: standard errors in parentheses

Table 6		
Effects of interest rates shocks on transition probabilities		
a/ $\xi_t = -100\text{b.p.}(t \text{ to } t+2)$		
Monthly		
	Before	After
	$p_{rr}$ 0.87	0.76
	$p_{re}$ 0.13	0.24
b/ $\xi_t = +100\text{b.p.}(t \text{ to } t+2)$		
	Before	After
	$p_{ee}$ 0.90	0.84
	$p_{er}$ 0.10	0.16