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## The single monetary policy and domestic macro-fundamentals: Evidence from Spain

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**The single monetary policy and domestic macro-fundamentals:  
Evidence from Spain**

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**Abstract**

We model pre-euro Spanish monetary policy and use our findings to assess the compatibility of the interest rates set by the ECB since 1999 with Spanish macro-fundamentals. We find that in the 1990s Spain implemented successfully a monetary strategy tailored to its own domestic fundamentals; and by abolishing it to join the euro she has paid a cost in the form of a sub-optimal monetary policy. Spain's experience suggests a cautious approach with regards to the timing of further EMU enlargement.

**Keywords:** Spain; ECB; monetary policy; domestic fundamentals; compatibility

**JEL classification:** C51, C52, E43, E58, F37

## 1. INTRODUCTION

In recent years a number of theoretical studies have addressed the question of optimal single monetary policy (SMP) in the European Economic and Monetary Union (EMU) under asymmetric national economic shocks, preferences and structures (see Dixit 2001, Aksoy et al. 2002, Benigno 2004, Matsen and Røisland 2005, Lombardo 2006, Hugh Hallett 2008 and Brissimis and Skotida 2008). These studies are motivated by the fear that asymmetries can cause intra-EMU frictions regarding the direction of the SMP; uncertainty about its effectiveness; and conflicts between the SMP and national fiscal policies leading, in extreme circumstances, to doubts about the sustainability of the EMU. The same considerations had previously motivated a rich empirical literature (see e.g. Bayoumi and Eichengreen, 1997) on European Optimum Currency Areas (OCA) and underpinned the EMU-accession criteria set by the Maastricht Treaty. In short, preoccupation with the policy implications of intra-EMU asymmetries is neither new nor surprising.

What is surprising is that the recent theoretical debate is not matched by corresponding empirical research.<sup>1</sup> Previous studies (see e.g. European Central Bank 2003, Benalal et al. 2006, Campa and Mínguez 2006, Arghyrou et al. 2010) have documented significant intra-EMU asymmetries in a number of areas and some divergence between the SMP and national fiscal policies (see Hugh Hallett and Lewis, 2008). Yet, with the exceptions of Hayo and Hofmann (2006) and Arghyrou (2008)<sup>2</sup> they do not provide estimates of any incompatibility between the fundamentals of individual countries and the SMP. Without such estimates it is difficult to assess the scope for intra-EMU conflicts and

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<sup>1</sup> Empirical research on European monetary policy since 1999 has mainly focused on modelling the preferences of the European Central Bank and comparing them with those of its national predecessors, mainly the Bundesbank (see, among others, Domenech et al. 2002 and Surico 2007).

<sup>2</sup> The former focus on Germany for which they find that the ECB interest rates might be too high, contributing to an economic slowdown. The latter examine Greece for which they conclude that the ECB policy has been too relaxed, resulting in unnecessarily high inflation.

national governments' motivation to influence the SMP through fiscal expansions. As a result, the debate on optimal SMP largely takes place within an empirical vacuum.

This paper aims to contribute towards partial bridging of this gap. We focus on Spain, the EMU's fourth largest economy producing in 2007 12% of the EMU's total output. By virtue of her size Spain may possess enough bargaining power to exercise significant leverage on the ECB both on her own as well as by leading coalitions of countries sharing its asymmetries against the EMU average (see e.g. Di Bartolomeo et al., 2006). Indeed, for Spain the question of compatibility had been raised even before her accession to the EMU in 1999 (see e.g. Gali, 1998). At the time there was plenty of room for optimism<sup>3</sup> as the 1990s had seen a significant improvement in Spanish macroeconomic performance reflected in all leading macro-indicators (see Figure 1). Since 1999 Spain has continued outgrowing the EMU's average and seeing its unemployment fall. However, its overall macroeconomic outlook has deteriorated. Most notably, Spanish inflation relative to the EMU has increased causing real effective exchange appreciation and negative real interest rates. These were followed by macroeconomic imbalances including a record-high current account deficit and dramatic price increases in the Spanish real estate market.<sup>4</sup>

The source of these adverse developments is debatable though fiscal policy is unlikely to be given that post-1999 Spain has turned its budget deficit into a surplus. One possibility is the existence of Balassa (1964) – Samuelson (1964) effects, an equilibrium-preserving mechanism preventing high-growth countries such as Spain from turning productivity gains into large current account surpluses at the expense of other EMU members. This argument, however, is rather unconvincing given the post-1999 low growth

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<sup>3</sup> Gali (1998) suggested the construction of an index of monetary tension to measure the difference between the interest rate produced by the ECB Taylor model and the BOS one. Carrying out a simulation study based on Taylor rules he concluded that the cost of abolishing monetary independence would be limited as long as Spain maintained business cycle synchronization with the euro zone.

<sup>4</sup> According to data provided by the BOS, between 1999 and 2007 house prices in Spain increased by 187%.

of Spanish productivity and record-high Spanish current account deficits which cannot be explained only by the income catch-up process (see Arghyrou and Chortareas, 2008).<sup>5</sup>

A second possibility is the existence of credibility gains caused by Spain's accession to the EMU (see Giavazzi and Spaventa, 1990). According to this hypothesis, by eliminating all intra-EMU currency risk euro-accession was bound to cause capital inflows, lower real interest rates, higher inflation and increased external deficits. In the medium-term, however, the resulting competitiveness losses are bound to lower demand and inflation so that real interest rates converge to the EMU average and the current account reverts to a sustainable position. These effects may well be in operation. However, if they were the main driving force of economic developments since 1999, ten years on the predicted medium-term reversion to a more sustainable equilibrium should be observed in the data. Instead, Spain's current account deteriorates fast and Spanish real interest rates do not show any evidence of convergence to the EMU average (see Arghyrou et al., 2010).

Finally, a third explanation, based on studies such as Honohan and Lane (2003) and Campa and Mínguez (2006), is that Spain's post-1999 macro-performance reflects deeper-seeded asymmetries in external trade links, financial structures and nominal and real rigidities causing heterogeneous transmission of the SMP to the Spanish economy relative to the EMU average. In other words, Spain's post-1999 macro-imbalances may be due to incompatibility between the SMP and Spain's fundamentals. This hypothesis has been adopted by numerous observers, including high-ranking Spanish policy-makers,<sup>6</sup> yet it has never been put to the test. It is precisely this assumed incompatibility we aim to test and quantify.

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<sup>5</sup> In relation to this, Rogers (2007) finds that Balassa-Samuelson effects on their own cannot fully account for persistent intra-EMU inflation differentials.

<sup>6</sup> For example, in 2006 when the ECB interest rate stood at 3%, the director of the Bank of Spain Research Division argued that Spain needed an interest rate of 4.5% (see Malo de Molina, 2006).

Our analysis follows a three-stage approach. First, we model Spanish monetary policy prior to euro-accession. Second, we forecast the interest rates the Bank of Spain (BOS) would have set after 1999 under a hypothetical regime of monetary independence. Finally, we use the difference between our forecasts and the actual ECB rates as a measure of compatibility between the SMP and Spanish fundamentals. Our findings can be summarised as follows. First, during 1980-98 Spanish monetary policy experienced four regime changes, all increasing the weight of inflation in the BOS's adopted policy rules. Second, during the last regime of Spanish monetary independence (1991-1998), the BOS was strongly inflation-averse and, despite its strong link to the German Mark, heavily influenced by domestic fundamentals. Third, after 1999 the BOS would have set interest rates twice as high as those set by the ECB. Overall, we conclude that Spain's accession to the EMU has come at a significant cost in the form of sub-optimal monetary policy. Our analysis has implications for the future enlargement of the EMU for whose timing it suggests a cautious approach.

The remainder of the paper is structured as follows: Section 2 reviews monetary developments in Spain over 1980-98. Section 3 presents our data, benchmark modelling methodology and tests for structural change in Spanish monetary policy during the pre-euro era (1980-98). Section 4 models formally Spanish monetary policy during the last regime of monetary independence (1991-98). Section 5 presents our compatibility analysis. Finally, section 6 offers concluding remarks.

## **2. PRE-EURO SPANISH MONETARY POLICY: A BRIEF DESCRIPTION**

The past three decades have seen a transformation of the Spanish financial system from a closed, strongly regulated one to an open, fully-liberalised market. Over the same period Spanish monetary policy underwent drastic changes in terms of its final goal,

intermediate targets and policy instruments (see Ayuso and Escrivá, 1997 and Gadea, 2000). More specifically, in the 1970s and up to 1983 the BOS followed money supply (M3) targets. These, however, became increasingly problematic due to money-demand instability caused by new methods of payment, inclusion of M3 in the liquid assets held by the public (LAP) and exclusion from the latter of the fast-increasing liquid short-run public debt<sup>7</sup> accompanying the fiscal expansion following the democratic transition of 1974.<sup>8</sup> In 1984 M3 was replaced by LAP as the intermediate targeted variable. This was subsequently relegated to a privileged monetary indicator as the BOS adopted a monetary policy converging to nominal GDP targets. This lasted until 1995 when the intermediate-targets framework was abolished and replaced by a strategy of direct inflation control.

With regards to the instruments of monetary policy, up to the mid-1980s the BOS targeted the volume of the Bank System Cash Assets (BSCA). Following the introduction of financial liberalisation BSCA was replaced by controls over short-run interest rates. In addition, following Spain's accession to the EU in 1986 the BOS adopted an implicit exchange rate target against the German mark. Within this framework the decision to join the Exchange Rate Mechanism (ERM) in 1989 pursued two inter-linked objectives. First, to establish a credible external anchor for Spanish monetary policy. Second, to reduce exchange rate speculation against the peseta. However, the BOS faced increasing difficulties in achieving these twin goals using a single policy instrument, the short-run interest rate. High profitability rates in the second half of the 1980s attracted large volumes of foreign capital causing increasing liquidity and peseta's appreciation. As a result, inflation pressures persisted prompting the BOS to resort to higher nominal interest rates.

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<sup>7</sup> The definition of the Bank of Spain for Liquid Assets held by the Public includes broad money supply (M3); liabilities of credit institutions and money-market instruments (such as insurance-like liabilities with savings banks); and treasury notes or bills held by the public.

<sup>8</sup> The strong influence that fiscal policy has historically exerted over monetary policy in Spain has been documented by Sabaté et al. (2006) and Gadea et al. (2008).



These, in turn, caused higher returns on public debt, further exchange rate appreciation and loss of external competitiveness. The paradoxical combination of weak macro-fundamentals and strong peseta rendered the latter vulnerable to speculative attacks during the ERM crisis of 1992-93 when it was devalued on three occasions.

Following the ERM reform in 1993 with the adoption of broad bands of fluctuation, Spanish monetary policy gradually relaxed. In 1994 the Law of BOS Independence, one of the requirements of the Maastricht Treaty, was passed by Parliament. This enhanced the credibility of Spanish monetary policy and laid the foundations for a successful convergence process based, since 1995, on direct inflation objectives. Following another devaluation of the peseta in 1995, Spanish interest rates commenced a systematic decline supported by fiscal consolidation and wage moderation. Consequently, and despite the high growth rates recorded in the second half of the 1990s, inflation pressures were contained. This allowed the peseta to converge smoothly towards its conversion parity against the euro which it joined in January 1999.

### **3. METHODOLOGY, DATA AND STRUCTURAL CHANGE ANALYSIS**

#### **3.1. Benchmark models and data definitions**

The benchmark specification we use to model Spain's pre-euro monetary policy is the forward-looking monetary policy reaction function proposed by Clarida et al. (1998). This models the short-term nominal interest rate  $r_t$  as a weighted average of a target level  $r^*_t$  and its own lagged value i.e.  $r_t = (1-\rho) r^*_t + \rho r_{t-1} + v_t$  where  $\rho$  denotes the degree of interest rate smoothing ( $0 \leq \rho < 1$ ) and  $v_t$  is a white noise policy shock term. The target interest rate is given by  $r^*_t = \bar{r} + \beta [E(\pi_{t+n}/\Omega_t) - \pi^*] + \gamma [E(y_t/\Omega_t) - y^*_t]$ , where  $\bar{r}$  denotes the long-run equilibrium nominal interest rate;  $E(\pi_{t+n}/\Omega_t)$  the inflation rate expected  $n$  periods ahead,  $\Omega_t$  the current information set;  $\pi^*$  the target inflation rate;  $y_t$  current output; and  $y^*_t$

the target (full-employment) level of output.<sup>9</sup> In this context, if  $\beta > 1$  ( $\beta < 1$ ), monetary policy is inflation-averse (inflation-accommodating); if  $\gamma > 0$  ( $\gamma = 0$ ) authorities are (are not) concerned with output stabilisation. By defining  $\alpha \equiv \bar{r} - \beta \pi^*$  and approximating  $E(\pi_{t+n}/\Omega_t)$  by  $\pi_{t+n}$ , the actual inflation rate between periods  $t$  and  $t+n$ , Clarida et al. obtain

$$r_t = (1-\rho) \alpha + \rho r_{t-1} + (1-\rho) \beta \pi_{t+n} + (1-\rho) \gamma (y - y^*)_t + \varepsilon_t \quad (1)$$

where  $\varepsilon_t$  is a white noise error term. This can be econometrically estimated as

$$r_t = \alpha + \beta_1 r_{t-1} + \beta_2 \pi_{t+n} + \beta_3 (y - y^*)_t + \varepsilon_t \quad (2)$$

from which  $r^*_t$  can be retrieved by setting  $\bar{r} = \alpha / (1 - \beta_1)$ ,  $\beta = \beta_2 / (1 - \beta_1)$  and  $\gamma = \beta_3 / (1 - \beta_1)$ .

Equations (1) and (2) can be extended to include a foreign interest rate to account for exchange rate targets such as those implemented by Spain in the 1990s.

We use quarterly data covering 1980:1-1998:4 taken from the ECB Databank provided by DataStream. We define  $r_t$  as the Spanish three-month inter-bank money market rate. The inflation rate  $\pi_t$  is given by  $\Delta_4 p_t$ , the percentage increase in Spanish CPI relative to the same quarter of the previous year;  $\pi_{t+n}$  is given by  $\Delta_4 p_{t+1}$ . Real output  $y_t$  is given by the Spanish seasonally-adjusted real GDP volume index (1995=100). Target output  $y_t^*$  is calculated fitting a Hodrick- Prescott (1997) filter into the  $y_t$  series.<sup>10</sup> Finally, we define the foreign interest rate to be the German day-to-day money rate,  $rger_t$ .

### 3.2. Unit root/stationarity tests

Estimating equation (2) using an appropriate framework presupposes knowledge of the variables' order of integration. We test for the latter using the unit root tests of Dickey and Fuller (1981), Phillips-Perron (1988), Perron and Ng (1998) and Ng and Perron

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<sup>9</sup> These are consistent with theoretical models of monetary policy where authorities have a quadratic loss function specified in terms of inflation and output. For further details see Clarida et al. (1998, p. 1037).

<sup>10</sup> We have also estimated  $y_t^*$  using the quadratic de-trending approach followed by Clarida et al. (1998). The resulting output gap series was similar to the one obtained by the Hodrick-Prescott filter, with the correlation coefficient between the two series being equal to 0.84.

(2001), as well as the KPSS test of stationarity of Kwiatkowski *et al.* (1992). The results are reported in Table 1, col. (a) to (d). With the exception of  $(y-y^*)_t$ , which is clearly stationary, the tests provide ambiguous conclusions. This is not surprising given the structural changes that occurred over 1980-1998, giving rise to potential biases towards rejecting stationarity (see Perron, 1989). Indeed, an eye-ball examination of Figure 2 suggests possible breaks in the mean and/or trend of the  $r_t$ ,  $\Delta p_t$  and  $rger_t$  series. We therefore apply the two-break Lagrange multiplier (LM) unit root test of Lee and Strazicich (2003), testing endogenously for changes in the series' intercept (Model A) and changes in the series' slope and intercept (Model C).<sup>11</sup> The results are reported in Table 1, col. (e). Using model C, we reject the null of a unit root at the 10 per cent level or lower.<sup>12</sup>

### 3.3. Testing for structural change in Spanish monetary policy

Having upheld stationarity for the variables in the benchmark model in equation (2) we test for structural change in Spanish monetary policy using the multivariate methodology of Bai and Perron (BP, 1998, 2003a, 2003b). This involves estimating a linear model allowing for  $m$  breaks, i.e.  $m+1$  regimes, in a relation of the form:

$$y_t = x_t' \beta + z_t' \delta_j + u_t \quad (3)$$

where  $y_t$  is the dependent variable;  $x_t$  ( $p \times 1$ ) and  $z_t$  ( $q \times 1$ ) are vectors of independent variables,  $\beta$  and  $\delta_j$  ( $j=1, \dots, m+1$ ) are the corresponding vectors of coefficients and  $T_1, \dots,$

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<sup>11</sup> In contrast to previously developed endogenous break unit root tests, the size properties of the minimum LM test are unaffected by breaks under the null. For further details see Lee and Strazicich (2003).

<sup>12</sup> Given our discussion in section 2, the breaks identified for  $r_t$ ,  $\Delta p_t$  and  $rger_t$  are consistent with our a priori expectations. Specifically, the breaks found for  $r_t$  coincide with increasing interest rates in 1988 (due to the BOS's effort to control inflation pressures following Spain's entrance into the ERM) and declining interest rates in the late 1990s as the latter started their convergence towards those of Germany in anticipation of EMU accession. The break found for  $\Delta p_t$  in 1988 reflects the onset of the previously discussed inflation pressures, whereas the one found in 1997 corresponds to the period during which the reduction in Spanish inflation is coming to a halt. Finally, the breaks identified for  $rger_t$  can be respectively linked to the tightening of German monetary policy at the end of the 1980s, aiming to cope with the inflationary effect of German reunification, and the monetary relaxation that took place in the mid-1990s.

$T_m$  are the break points treated endogenously by the model. To identify breaks BP propose three tests. First, the  $supF\{T\}(k)$  test, testing the null of no breaks against the alternative of  $k$  breaks. Second, the  $supF\{T\}(l+1/l)$  test, testing the null of  $l$  breaks, with  $l = 0,1,\dots,n$  against the alternative of  $l+1$  breaks. Finally, the “double maximum” tests,  $UDmax$  and  $WDmax$  testing the null of no structural breaks against the alternative of an unknown number of breaks. BP suggest starting the testing process with the sequential test  $supF\{T\}(l+1/l)$ . If no break is detected, they suggest applying the  $UDmax$  and  $WDmax$  tests to determine whether at least one break exists. If the null is rejected, they suggest continuing with a sequential application of the  $supF\{T\}(l+1/l)$  test.

We implement BP’s testing approach on a pure changing model of Spanish monetary policy given by equation (4) below:

$$r_t = \alpha + \beta_1 r_{t-1} + \beta_2 \Delta_4 p_t + \beta_3 \Delta_4 p_{t+1} + \beta_4 (y - y^*)_t + \beta_5 (y - y^*)_{t+1} + \beta_6 rger_t + \beta_7 rger_{t+1} + \varepsilon_t \quad (4)$$

Equation (4) is an extended version of the benchmark model in equation (2) allowing for foreign interest rate effects and including the current and one-period lead values of the model’s independent variables. A maximum number of 5 breaks has been considered, which, given the sample size  $T=110$  implies a trimming parameter 0.10. The process is allowed to present autocorrelation and heteroskedasticity and uses a non-parametric correction to take account of these effects. Table 2 presents the results. Three structural breaks are identified. They are located, with narrow confidence intervals, in 1984:1, 1987:1 and 1990:4. Given our discussion in Section 2, these are entirely consistent with our expectations. The first two breaks correspond to the gradual abandoning of BSCA controls in favour of short-run interest rates in 1984; and the adoption of an implicit exchange rate target in 1987 before the peseta joined the ERM. On the other hand, the break in 1990.4 reflects the beginning of the formal convergence process towards EMU

participation.<sup>13</sup> Overall, our findings suggest that during the period under consideration Spanish monetary policy includes four regimes, respectively covering 1980-83, 1984-86, 1987-90 and 1991-98.<sup>14</sup>

The small number of available quarterly observations does not allow an individual econometric investigation of the first three regimes.<sup>15</sup> Nevertheless, we can still obtain qualitative indications about the effects these breaks have caused by estimating a restricted version of equation (4), excluding the variable's first lead values, for the whole of the period 1980-98 and examining the model's recursively estimated parameters and  $t$ -values.<sup>16</sup> These are reported in Figure 3, revealing an increasing weight attached to inflation and a simultaneous reduction in the weight of output gap. The latter, however, remains statistically significant throughout most of our sample with its significance increasing substantially after 1992. Furthermore, we obtain a declining constant parameter, indicative of a declining risk premium in Spanish interest rates, i.e. increasing credibility for Spanish monetary policy. Therefore, and quite interestingly, Figure 3 suggests an increase in the significance of domestic fundamentals in the 1990s. Over the same period, we also observe a rapidly increasing role for the German interest rate, which is consistent

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<sup>13</sup> In addition to equation (4), we have also estimated two partial changing models respectively accounting only for breaks in the inflation and output gap parameters. For the model allowing the inflation parameters to change we obtain exactly the same three breaks obtained by the pure changing model. For the model allowing changes in the output gap parameters, we find one structural break in 1984:4. The results are available upon request.

<sup>14</sup> The literature estimating monetary policy rules for the BOS is scarce and, to the best of our knowledge, has not derived structural breaks endogenously as we do in this paper. López (2002) analyses Spanish monetary policy from 1984 to 1998. He introduces a single, exogenous structural break in 1993 and confirms that the BOS followed a Taylor-type interest rate rule. Díaz and Montero (2004) study Spanish monetary policy in the context of two exogenously chosen periods, 1978-1989 and 1989-1998. They also find that over the second period the monetary policy of the BOS is well characterized by a Taylor-type rule.

<sup>15</sup> We have attempted individual econometric investigation of each of the first three sub-periods using monthly data without obtaining any statistically significant results.

<sup>16</sup> More specifically our estimated model is  $r_t = \alpha + \beta_1 \Delta p_{t-1} + \beta_2 (y-y^*)_{t-1} + \beta_3 rger_t + u_t$ . The results do not change if instead of the first lag of  $\Delta p_t$  and  $(y-y^*)_t$ , we use the current values of these variables. However, first lags were preferred to avoid simultaneity problems with  $r_t$ . We could have accounted for endogeneity by estimating the static version of the model using instrumental variables. However, as Spanish monetary policy has been subject to multiple structural breaks, different instruments might be suitable in different periods. Finally, note that we have used the current value rather than the first lag of  $rger_t$ , as we can plausibly assume that during the period under investigation Spanish monetary policy was exogenous to the German one.

with the Spanish exchange rate targets followed since 1987. The recursively estimated  $t$ -score of  $rger$ , however, is not significant at any point in time. This is not consistent with our expectations and merits further investigation, offered in Section 4 below.<sup>17</sup>

## 4. MODELLING PRE-EURO SPANISH MONETARY POLICY, 1991-1998

### 4.1. Linear models of monetary policy

This section focuses on the last regime of independent Spanish monetary policy, 1991-98, for which the number of available observations allows a detailed econometric investigation. We first estimate the benchmark model in equation (2). To avoid simultaneity problems we use the two-stage instrumental variables method.<sup>18</sup> We report our findings in Table 3 col. (a) and their long-run counterparts in Table 4, col. (a).<sup>19</sup> We obtain an average degree of interest rate smoothing and a coefficient of  $\Delta_4 p_t$  in the long-run target reaction function significantly greater than unity, thus suggesting non-accommodating monetary policy. However, contrary to our expectations and the evidence presented in Figure 3, the output gap coefficient is not significant. Adding  $rger_t$  to the equation results in very similar findings (see col. (b) in Tables 1 and 2). Furthermore, the coefficient of  $rger$  is not significant. Clearly, and given the strong link of Spanish to German monetary policy in the 1990s, this finding is not plausible.<sup>20</sup>

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<sup>17</sup> For robustness we have repeated this recursive estimation using data of monthly frequency and industrial production as a proxy for real GDP. The results (available upon request) remain unchanged.

<sup>18</sup> The instruments used are the first lag of inflation, the first lag of the output gap and the fifth lag of the German output gap. The Instrumental Variables (IV) method is similar to the Generalised Maximum Likelihood (GMM) method and has the advantage of being able to handle endogeneity among regressors using limited information, especially in behavioural models such as those we estimate (see Hendry and Doornik 2001, p. 167). Like the IV, the GMM method is instruments-based but more suitable when the form of the data density is known and the sample size is sufficiently large (see Hendry and Doornik 2001, p. 183). Given the relatively small number of observations available for our analysis and the large number of parameters to be estimated (especially in the case of the non-linear model presented below), the two-stage IV method was preferred.

<sup>19</sup> Note that the data favoured the use of the current inflation rate ( $\Delta p_t$ ) rather than its first lead.

<sup>20</sup> For robustness we have repeated these estimations using the series of monthly industrial production as a proxy for real GDP. The results (available upon request) remain unchanged.

## 4.2. Non-linear models of monetary policy

### 4.2.1. Non-linearity tests

We now seek to improve upon the linear models discussed above by testing for output gap effects in Spanish monetary policy, as predicted by theoretical models such as Orphanides and Wieland (2000) and empirically found for other countries (see e.g. Bec et al., 2002, Surico 2007 and Martin and Milas, 2004). To that end we first test for non-linear interest rate adjustment using the approach proposed by Teräsvirta and Anderson (1992), Granger and Teräsvirta (1993) and Teräsvirta (1994). This is based on equation (5) below:

$$r_t = \gamma_{00} + \sum_{j=1}^{\phi} (\gamma_{0j} r_{t-j} + \gamma_{1j} r_{t-j} r_{t-d} + \gamma_{2j} r_{t-j} r_{t-d}^2 + \gamma_{3j} r_{t-j} r_{t-d}^3) + \gamma_4 r_{t-d}^2 + \gamma_5 r_{t-d}^3 + v_t \quad (5)$$

Linearity is described by  $H_0: [\gamma_{1j} = \gamma_{2j} = \gamma_{3j} = \gamma_4 = \gamma_5 = 0]$  for all  $j \in (1, 2, \dots, \phi)$ , where  $d$  is a delay parameter,  $\phi$  is determined through inspection of the partial autocorrelation function of  $r_t$  and  $v_t \sim niid(0, \sigma^2)$ .<sup>21</sup>  $H_0$  can be tested using a general LM-type test, denoted by  $LM^G$ , estimated for all plausible values of  $d$ . Linearity is rejected if any of the  $LM^G$  statistics is significant. In that case, the optimum value of  $d$  is given by the highest  $LM^G$  score. To determine the specific type of non-linearity we follow the selection procedure proposed by Teräsvirta and Anderson (1992). We first test the null [ $H_0: \gamma_{3j} = \gamma_5 = 0; j \in (1, 2, \dots, \phi)$ ]. We denote this test by  $LM^{L1}$ . If significant,  $LM^{L1}$  implies logistic non-linearity. If  $LM^{L1}$  is insignificant, we test the null [ $H_0: \gamma_{1j} = \gamma_{2j} = \gamma_4 = 0 \mid \gamma_{3j} = \gamma_5 = 0; j \in (1, 2, \dots, \phi)$ ]. We denote this test by  $LM^Q$ . A significant  $LM^Q$  score implies quadratic non-linearity. If both  $LM^{L1}$  and  $LM^Q$  are insignificant, we calculate a third statistic,  $LM^{L2}$ , testing the null [ $H_0: \gamma_{1j} = 0 \mid \gamma_{2j} = \gamma_{3j} = \gamma_4 = \gamma_5 = 0; j \in (1, 2, \dots, \phi)$ ]. A significant  $LM^{L2}$  score implies logistic

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<sup>21</sup> Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing  $\phi$  using information criteria such as the Akaike, since this may induce a downward bias.

non-linearity. If  $LM^{L1}$ ,  $LM^{L2}$  and  $LM^Q$  are all significant, the type of non-linearity is determined by the strongest rejection of the null (see Teräsvirta, 1994, p.212).

Our non-linearity tests are reported in Table 5. The Partial Autocorrelation Function of  $r_t$  (not presented due to space constraints) suggested  $\phi = 1$ . Given the quarterly frequency of our data we set  $d = 1 \dots 4$ . For  $d=2$  the  $LM^G$  and  $LM^{L1}$  scores are significant at the 1 and 5 per cent level respectively, suggesting non-linearity of the logistic type. For  $d=2$   $LM^Q$  is also significant, however the rejection of the null by  $LM^{L1}$  is stronger, confirming the existence of logistic non-linearity.<sup>22</sup>

#### 4.2.2. A model of monetary policy with output gap effects

We now model the logistic non-linearity found in Spanish nominal interest rate using the Logistic Smooth Threshold Error Correction Model (L-STECM).<sup>23</sup> This is given by equations (6) to (9) below, where  $\varepsilon_t$ ,  $u_{1t}$  and  $u_{2t}$  are white noise error terms:

$$r_t = \theta_t r_{1t} + (1 - \theta_t) r_{2t} + \varepsilon_t \quad (6)$$

$$r_{1t} = \alpha_1 + \beta_{11} r_{t-1} + \beta_{12} \Delta_4 p_t + \beta_{13} (y - y^*)_t + \beta_{14} rger_t + u_{1t} \quad (7)$$

$$r_{2t} = \alpha_2 + \beta_{21} r_{t-1} + \beta_{22} \Delta_4 p_t + \beta_{23} (y - y^*)_t + \beta_{24} rger_t + u_{2t} \quad (8)$$

$$\theta_t = pr \{ \tau \geq (y - y^*)_{t-d} \} = 1 - \frac{1}{1 + e^{-\sigma[(y - y^*)_{t-d} - \tau]}} \quad (9)$$

Equation (6) models the nominal interest rate as a weighted average of two regimes, a lower ( $r_{1t}$ ) and an upper ( $r_{2t}$ ), respectively corresponding to periods of normal and non-normal (overheating) output conditions. The regime applying each period is determined according to whether the model's transition variable takes values below or

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<sup>22</sup> For robustness we have repeated our non-linearity tests using data of monthly frequency. The results (available upon request) confirmed the existence of non-linearity of logistic type with a delay parameter equal to four months. This is consistent with the delay parameter of two quarters found above.

<sup>23</sup> This model has been used to model logistic non-linearities in UK monetary policy by Martin and Milas (2004). A thorough discussion of the L-STECM and other non-linear, smooth-adjustment error correction models can be found in van Dijk et al. (2002).



above a critical threshold  $\tau$ . Given our previous finding of  $d=2$ , we define our transition variable to be  $(y-y^*)_{t-2}$ . The weight attached to the “normal” regime in equation (6),  $\theta_t$ , is given by equation (9) as the probability that the transition variable takes values below  $\tau$ , where the parameter  $\sigma$  is the speed of transition between the two regimes.<sup>24</sup>

The parsimonious estimate of the L-STEEM, obtained after eliminating all statistically insignificant terms, is reported in Table 3, col. (c) with its long-run counterpart in Table 4, col. (c). The econometric properties of the L-STEEM are superior to its linear counterparts as it passes all misspecification tests and produces a significantly lower regression standard error. For the lower regime, we obtain an average degree of interest rate smoothing and, consistent with our expectations, a statistically significant positive coefficient for  $(y-y^*)_t$  and  $rger_t$ . In the upper regime the sole determinant of Spanish monetary policy is  $\Delta p_t$  with a long-run coefficient significantly higher than unity.<sup>25</sup>

Overall, the findings of our preferred L-STEEM specification confirm the significant effect of German interest rates on Spanish monetary policy during the 1990s. At the same time they suggest that domestic fundamentals also played a very significant role in determining Spanish interest rates. In other words, in the 1990s the BOS did not simply shadow the mark; rather it used its exchange rate targets as only one of the components formulating a successful domestically-oriented monetary strategy tailored towards Spanish fundamentals. The latter’s abolition in 1999 implies that Spain’s accession to the euro may have come at a significant cost, an issue to which we now turn our attention.

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<sup>24</sup> In practice  $\sigma$  is usually estimated very imprecisely as the likelihood function in (7) is very insensitive to this parameter (see the detailed discussion on this point in van Dijk et al., 2002).

<sup>25</sup> For robustness we have estimated the L-STEEM model using data of monthly frequency and industrial production as a proxy of real GDP. The results (available upon request) were consistent with those obtained using quarterly data.

## 5. THE ECB AND SPANISH MACRO-FUNDAMENTALS

In this section we assess the compatibility between the SMP and domestic Spanish fundamentals using a counterfactual experiment. This consists of comparing the actual interest rates set by the ECB against the out-of-sample forecasts obtained for the Spanish target interest rate from the equations reported in Table 4, given the values of Spanish macro-fundamentals during 1999:1-2007:2 and assuming that the BOS would have maintained its policy preferences of the 1990s.<sup>26</sup> The results are presented in Figure 4(a) and Table 6. All models provide average forecasts for the Spanish interest rates approximately double than the actual ECB ones. This implies that post-1999 the BOS would have followed a much tighter monetary policy than the ECB.

A caveat relating to these findings is that they do not account for a possible reduction in Spain's equilibrium real interest rate due to the credibility gains predicted by Giavazzi and Spaventa (1990). Hayo and Hoffman (2006) and Arghyrou (2008) address this point by adjusting the constant in equation (2) for a lower equilibrium EMU real interest rate ( $\bar{r}_{EMU}$ ) and a target inflation rate equal to the ECB's inflation objective of 2% ( $\pi^*_{ECB}$ ). The adjusted constant is given by  $\alpha^{adj} \equiv \bar{r}_{EMU} + (1-\beta) \pi^*_{ECB}$  where  $\beta$  is the estimated coefficient of inflation in the pre-euro monetary policy target reaction function. Following Clarida et al. (1998, p.1046), we set  $\bar{r}_{EMU}$  equal to the average value of the ex-post EMU-average real interest rate for the period 1999-2007, i.e. 1.15%, which is very close to the 1.28% formally derived by Hayo and Hoffman. The results are presented in

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<sup>26</sup> This approach, also used by Hayo and Hoffman (2006) and Arghyrou (2008) for Germany and Greece respectively, is based on a number of assumptions regarding the state of the world that would have been prevailed had Spain not joined the EMU in 1999. These are, by definition, subject to questioning. However, it is reasonable to assume that had Spain not joined the euro, the BOS would have continued to set Spanish monetary policy taking into account Spanish domestic output and inflation conditions. In addition, as Clarida et al. (1998, p. 1058) highlight for their own analysis, our counterfactual analysis does not capture the likely effects EMU accession may have caused on Spanish inflation and output. This would involve specifying and estimating a complete macroeconomic model for the Spanish economy, which exceeds the scope of the present paper. However, our analysis below accounts for the possibility of a lower equilibrium real interest rate caused by Spain's accession to the euro.

Table 6 and Figure 4(b). Compared to the initial projections, our findings do not change substantially, although the L-STEEM suggests a lower degree of incompatibility over the period 2003-2005. Nevertheless, this seems to have re-emerged since early 2005.<sup>27</sup>

Overall, our analysis suggests significant incompatibility between the SMP and Spanish macro-fundamentals. The intuition underlying this finding is that in view of the high Spanish growth rates of 1999-2007, given its strong inflation aversion in the 1990s, the BOS would have set higher nominal interest rates to control inflation pressures. By contrast, the ECB followed a policy of low interest rates fuelling Spanish demand further and allowing inflation pressures to cause higher actual relative inflation. As argued earlier, the latter have caused increasingly negative real interest rates and significant appreciation in Spain's real effective exchange rate. These, in turn, have resulted in significant macroeconomic imbalances including a bubble in the real estate market (see e.g. Fernandez-Kranz and Hon, 2006) and record current account deficits not explained by the income catch-up process (see Arghyrou and Chortareas, 2008). As a result, our findings suggest that by abolishing its successful monetary to join the euro in 1999, Spain has paid a significant accession cost in the form of a sub-optimal monetary policy.

## **6. CONCLUDING REMARKS**

In recent years a number of studies have focused on optimal single monetary policy (SMP) in the European Economic and Monetary Union (EMU) under asymmetric national economic shocks, preferences and structures. This theoretical literature has not been accompanied by an equal volume of empirical research. In this paper we model pre-euro Spanish monetary policy and use our findings to assess the compatibility of the SMP with Spanish macro-fundamentals. We find that by the 1990s Spain had established an inflation-

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<sup>27</sup>For this model the adjusted constant is calculated using the coefficient of inflation in the upper regime, as in the lower regime inflation was not found to be statistically significant.

averse monetary framework tailored to its own requirements and delivering optimal monetary management; and by abolishing it in 1999 to join the euro Spain has paid a significant cost in the form of a SMP incompatible with its domestic fundamentals.

Our findings contribute towards establishing a critical mass of evidence indicating that the SMP, institutionally oriented towards the EMU average, may be contributing to divergent intra-EMU national business cycles. If true, this has two implications. First, despite achieving nominal convergence in the 1990s, the EMU remains a non-optimum currency area. Second, the concerns motivating theoretical research on optimal monetary policy rules, namely that ECB may be put under conflicting pressures from individual EMU members or coalitions of members, are not unfounded. To avoid this risk, structural reforms promoting real business cycles convergence in the EMU area appear necessary.

This observation has implications for the future enlargement of the EMU. If all countries that joined the EU in 2004 were at present EMU members, the group of members outperforming the average union's growth rate would increase its share in the union's total output from its current 15%<sup>28</sup> to 25% and account for 40% of the union's population. This might tilt the SMP towards higher interest rates, not necessarily welcomed by the EMU's core countries. In that case the scope for internal tensions and monetary policy uncertainty could increase significantly. Therefore, it might be appropriate for the new EU countries to join the euro after having achieved a higher degree of real convergence than countries such as Spain and Greece had done by the time of their own accession. Recent research suggests that such convergence is indeed taking place.<sup>29</sup> Nevertheless, until it is consolidated and evident in a wide range of areas, it might be optimal for all parties involved to adopt a cautious approach with respect to the timing of further EMU enlargement.

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<sup>28</sup> This figure includes Cyprus, Greece, Ireland, Spain and Slovenia.

<sup>29</sup> See, among others, Boeri and Galibardi (2006), Eickmeier and Breitung (2006) and Arghyrou et al. (2010).

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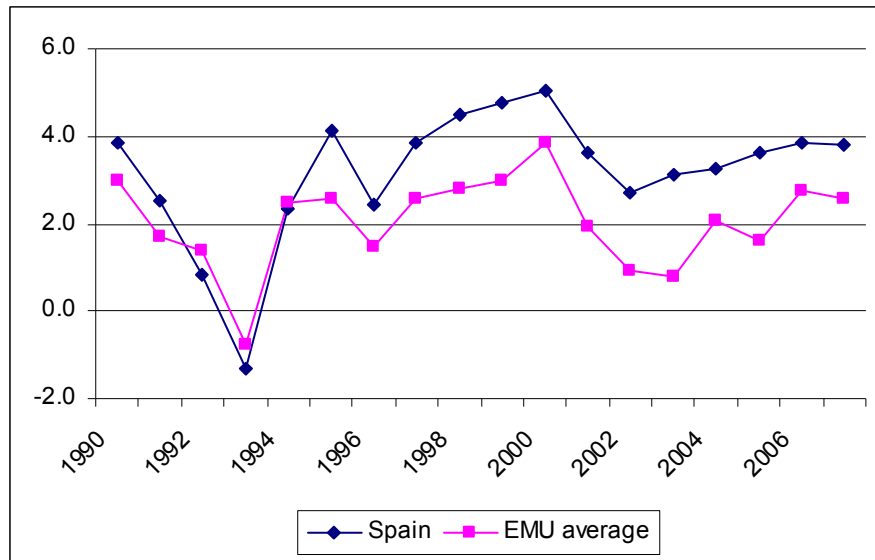
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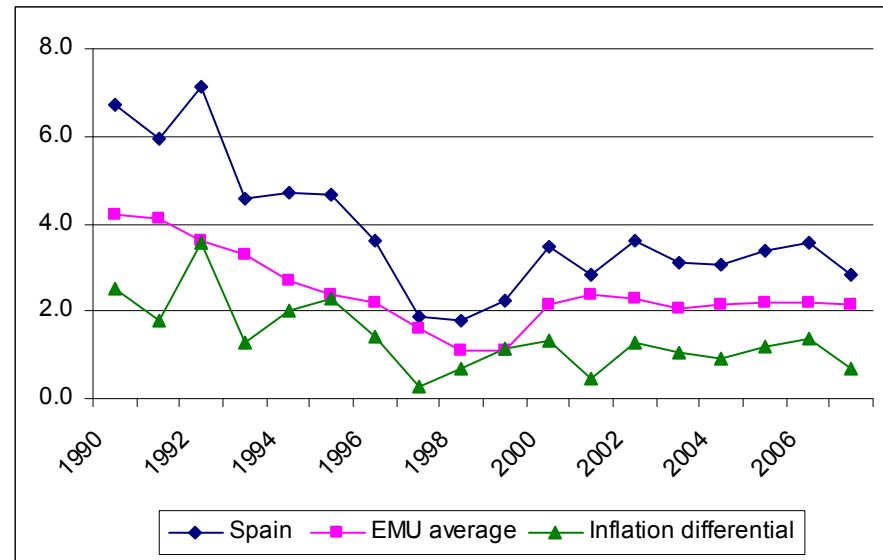
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Figure 1: Leading macroeconomic indications in Spain and the EMU, 1990-2007

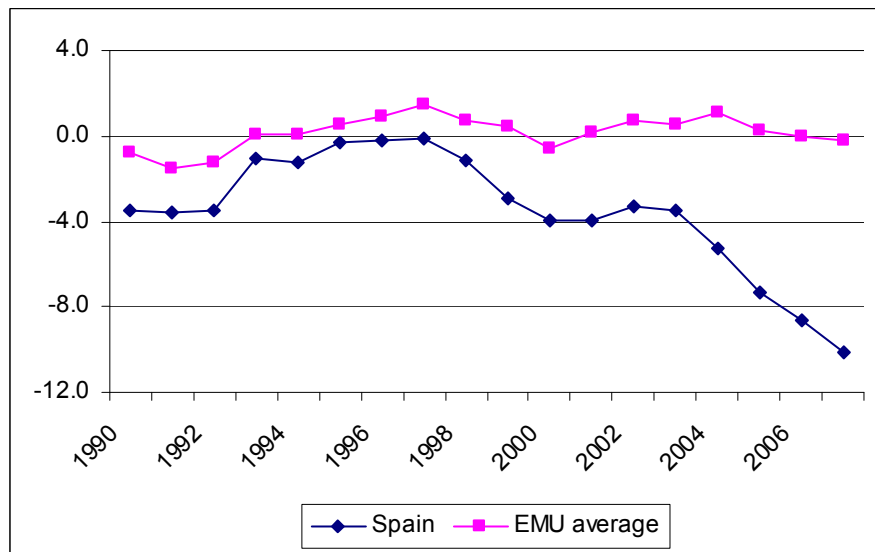
(a) GDP growth (%)



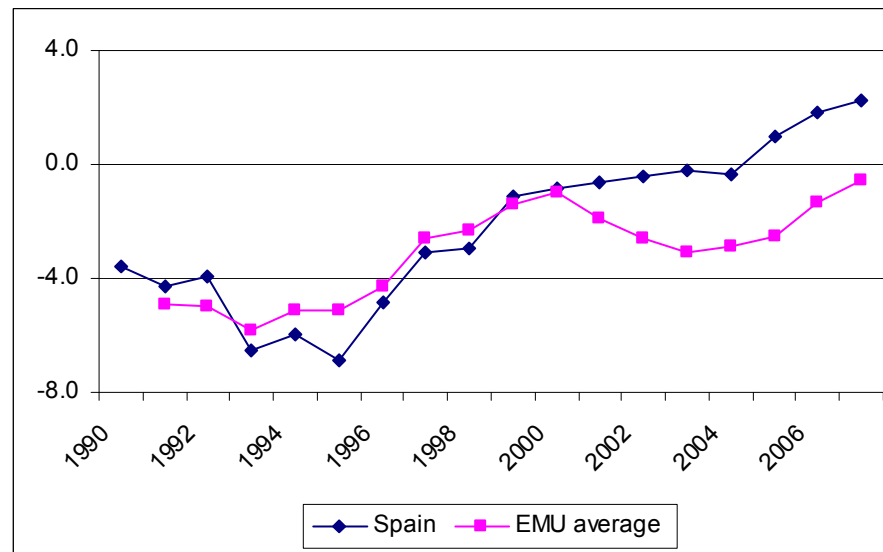
(b) CPI inflation (%)



(c) Current account balance (% of GDP)



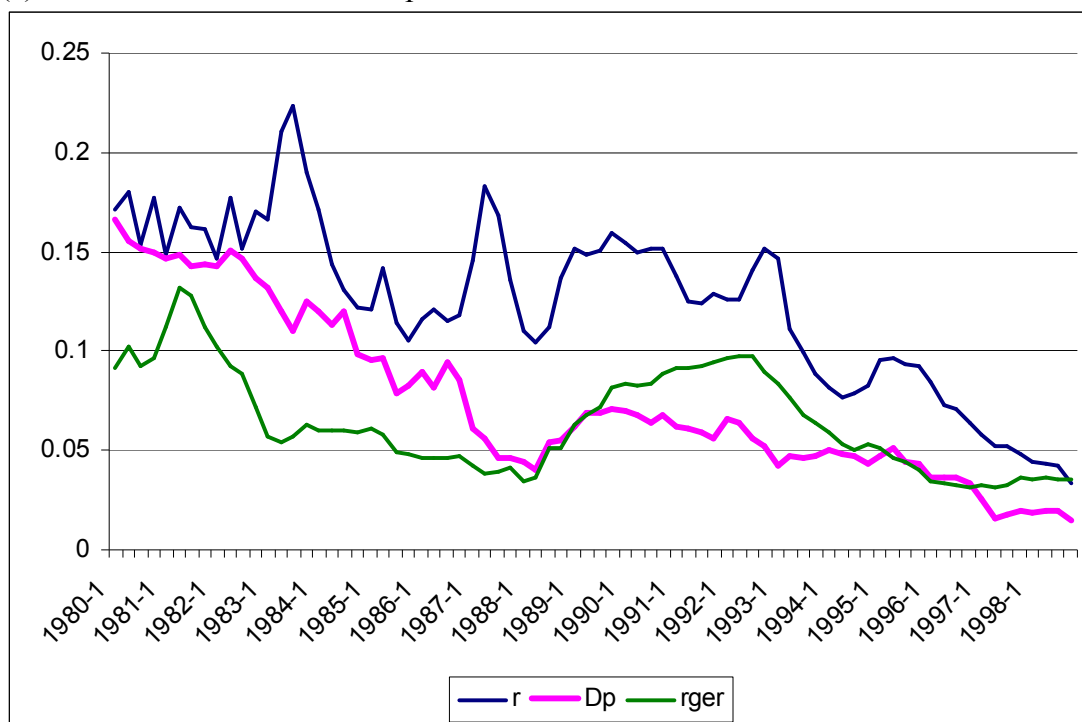
(d) Public budget balance (Maastricht definition, %)



Source: International Financial Statistics

Figure 2: Data description, 1980:1-1998:4

(a) Nominal interest rates and Spanish CPI inflation



(b) Spanish output gap

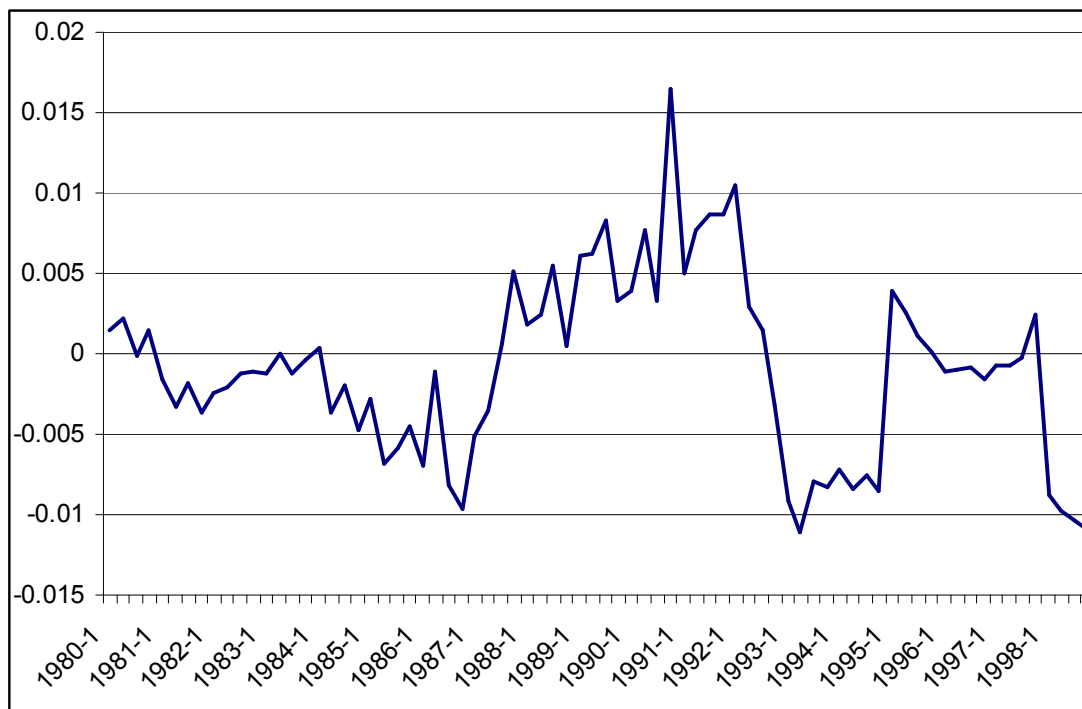
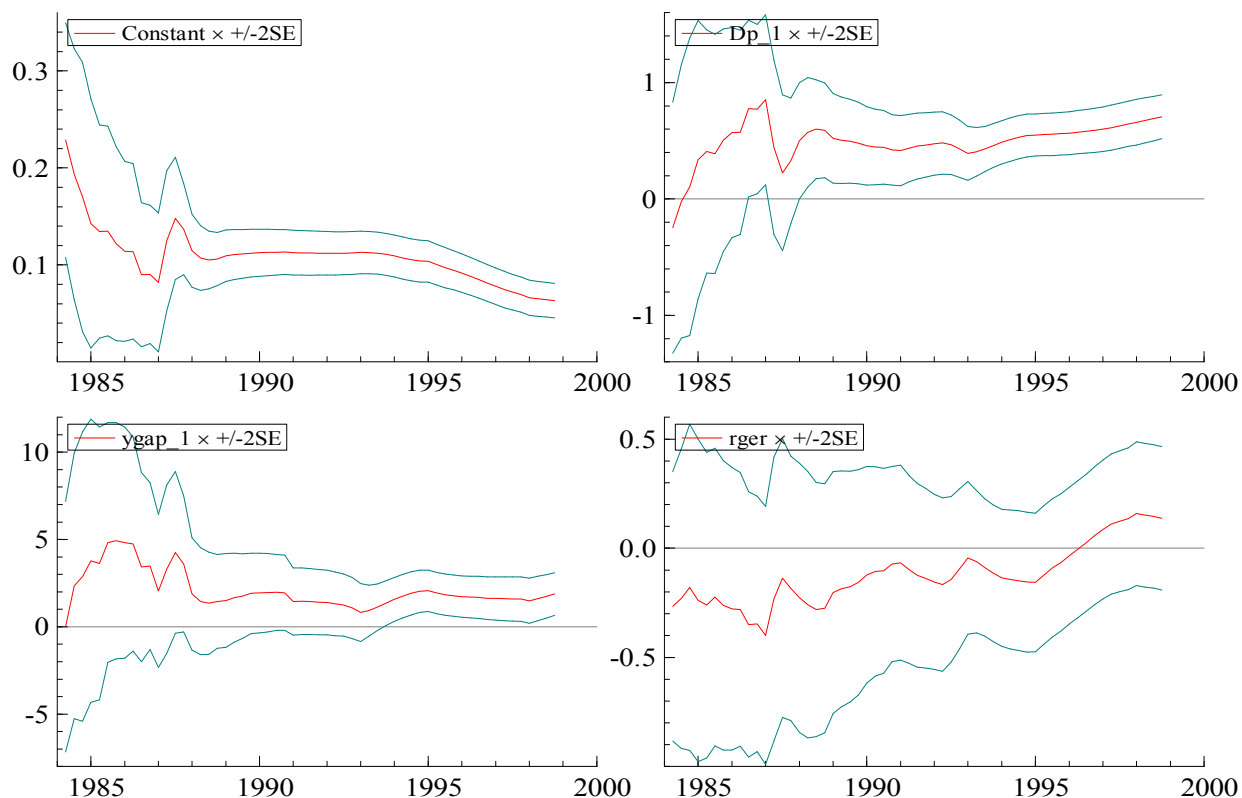


Figure 3: Recursive estimates of pure changing model

(a) Recursive parameter coefficients  $\pm 2 \times$  standard errors



(b) Recursive t-scores

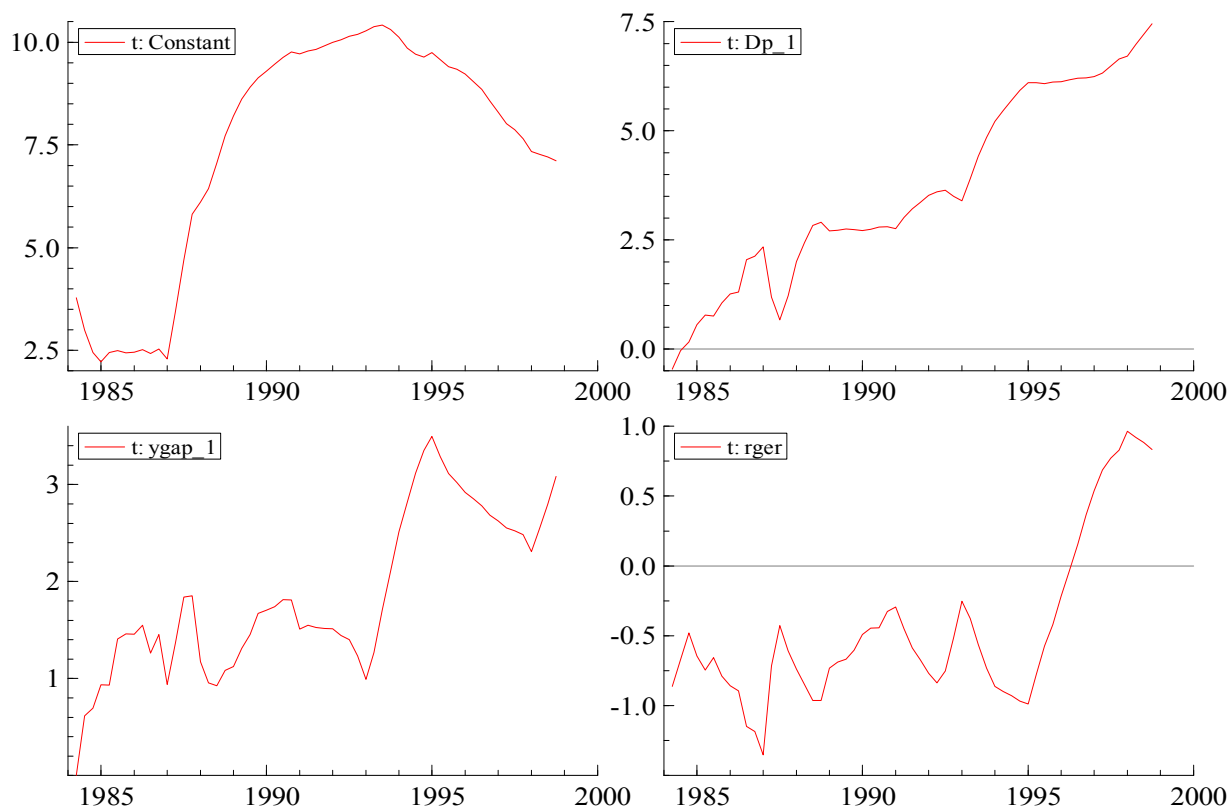
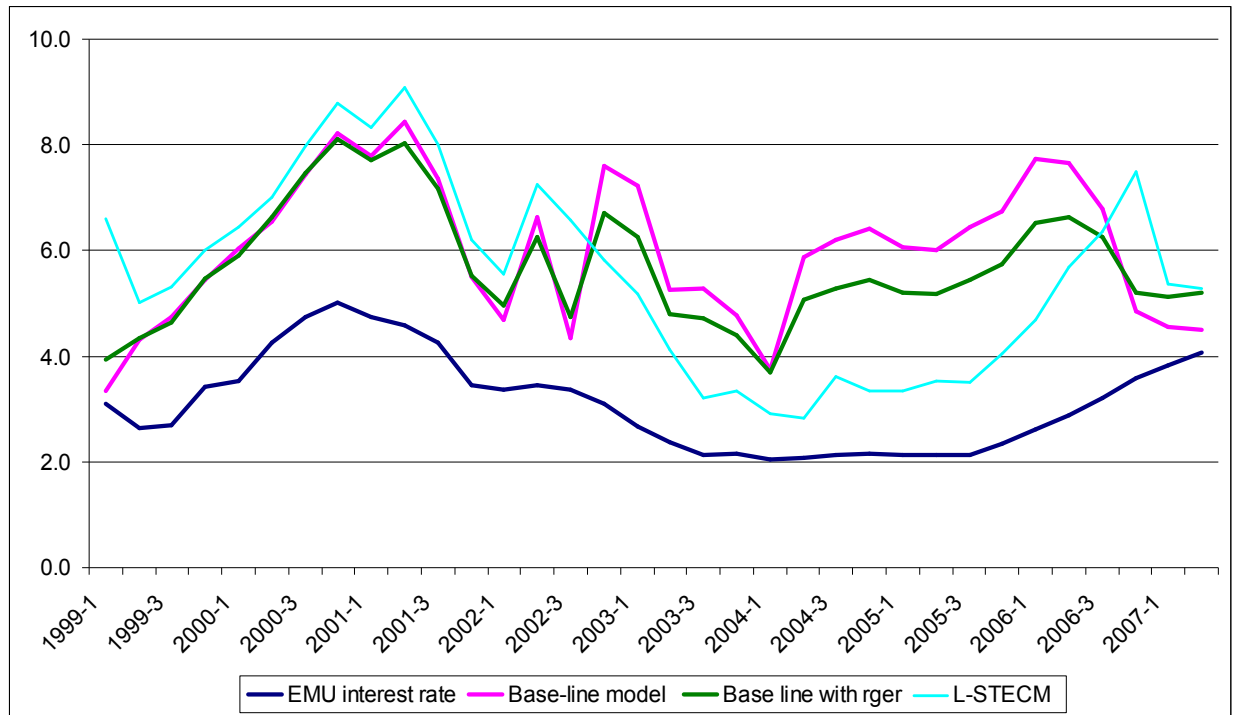


Figure 4: Actual EMU versus projected Spanish nominal interest rates, 1999:1-2007:2

(a) Projections with non-adjusted constant



(b) Projections with adjusted constant



Table 1: Unit root/stationarity tests

	(a)	(b)	(c)	(d)		(e)
With constant and trend	ADF	PP	MZ <sub>t</sub> -GLS	KPSS ( $\eta_\tau$ )	LS – Model (A)	Break points
$r_t$	-1.44	-1.23	0.69	1.11**	-3.22	1985:1, 1993:1
$\Delta_4 p_t$	-1.78	-3.39*	0.77	0.97**	-1.80	1984:3, 2002:3
$(y-y^*)_t$	-3.75*	-3.67*	-3.27**	0.06	-4.53*	1991:3, 1995:3
$rger_t$	-2.44	-1.88	-1.15	0.69*	-4.94**	1983:2, 1992:4
With constant	ADF	PP	MZ <sub>t</sub> -GLS	KPSS ( $\eta_\tau$ )	LS – Model (C)	Break points
$r_t$	-3.38+	-3.32+	-3.30*	0.11	-5.01+	1988:3, 1996:3
$\Delta_4 p_t$	-1.78	-2.17	-0.77	0.27**	-5.12+	1988:2, 1997:2
$(y-y^*)_t$	-3.76*	-3.68*	-3.31*	0.05	-5.37+	1986:3, 1992:2
$rger_t$	-2.58	-2.20	-2.49	0.07	-7.09**	1988:3, 1994:4

Notes: \*\*, \*, + respectively denote significance at the 1, 5 and 10 per cent level. The critical values of the ADF and PP test can be found in McKinnon (1996). The critical values of the two-break LM unit root test can be found in Lee and Strazicich (2003). The number of lags of in ADF tests has been selected in accordance with Ng and Perron (1995); those of the MZ<sub>t</sub>-GLS test have been determined by the Schwartz Bayesian Information Criterion. For the PP test, Bartlett's window has been used as a kernel estimator, choosing the bandwidth in the PP and KPSS test by the Newey and West (1994) method.

Table 2: Testing for structural breaks in Spanish monetary policy

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SupF <sub>T</sub> → no breaks $m=0$ versus $m=k$ breaks:				
$k=1$	$k=2$	$k=3$	$k=4$	$k=5$
71.82**	105.33**	79.83**	54.48**	40.93**
No breaks versus undetermined number of breaks:				
UDmax	105.33**			
WDmax	121.07**			
SupF <sub>T</sub> ( $l+1/l$ ) → $l$ breaks versus $l+1$ breaks:				
$l=1$	$l=2$	$l=3$	$l=4$	
85.02**	45.48**	7.24	0.00	
Number of breaks selected Sequential method:			3	
Break point and confidence intervals				
$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$		
1984.1	1987.1	1990.4		
(1983.4, 1984.2)	(1986.3, 1987.2)	(1990.2, 1991.1)		

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Notes: \*\*, denotes significance at the 1 per cent %; p-values in brackets. Critical values are taken in Bai and Perron (1998). Changes in the pure structural model given by equation (4) are tested selecting a trimming parameter  $\varepsilon=0.10$  and a maximum number of 5 structural breaks. Serial correlation in the errors is not allowed. The consistent covariance matrix is constructed using the Andrews (1991) method.

Table 3. Monetary policy reaction functions

Variable	(a)	(b)	(c)	
	Linear model	Linear model with <i>rger</i>	Non-linear L-STECM model	
			Lower regime	Upper regime
Constant	-0.002 (0.005)	-0.001 (0.006)	0.001 (0.003)	0.001 (0.003)
$r_{t-1}$	0.733 (0.097)**	0.641 (0.109)*	0.655 (0.071)**	
$\Delta_4 p_t$	0.565 (0.239)*	0.491 (0.243)*		2.174 (0.051)**
$(y-y^*)_t$	0.204 (0.340)	0.169 (0.331)	1.818 (0.330)**	
<i>rger</i> <sub>t</sub>		0.191 (0.139)	0.648 (0.143)**	
$\tau$			0.001 (0.0005)*	
Instrument Val.	0.11	0.16	N/A	
Reg. S.E.	0.00857	0.00831	0.00397	
AR	0.13	0.13	0.10	
ARCH	0.02	0.01	0.77	
Norm	0.11	0.17	0.82	
Hetero	0.00	0.00	0.87	

\*\* , \* denotes significant at the 1 and 5 per cent level. Numbers in parentheses are standard errors, p-values in square brackets. Reg. S.E. stands for Regression Standard Error. AR is the Lagrange Multiplier F-test for third-order residual serial correlation; ARCH is the Autoregressive Conditional Heteroscedasticity F-test; Norm is the Normality Chi-square Bera-Jarque test for residuals' non-normality; Hetero is an F-test for heteroscedasticity; Instrument Val. is the Sargant Chi-square test for instruments' validity. The output of the econometric programme used for estimating non-linear models (Pc-Give) does not report instrument validity tests in non-linear algorithms.



Table 4 – Target monetary policy reaction functions

	(a)	(b)	(c)	
	Linear model	Linear model with $rger_t$	Non-linear Lower regime	L-STEMM Upper regime
Constant	-0.007	-0.003	0.003	0.001
$\Delta p_t$	2.116	1.367		2.174
$(v-y^*)_t$	0.765	0.470	5.270	
$rger_t$		0.531	1.875	

Table 5: Linearity tests on Spanish three-month money market rates

$d$	$LM^G$	$LM^{L1}$	$LM^Q$	$LM^{L2}$
1	0.81 [0.498]	N/A	N/A	N/A
2	5.56** [0.000]	4.76* [0.017]	3.53* [0.029]	0.81 [0.378]
3	1.18 [0.386]	N/A	N/A	N/A
4	0.42 [0.832]	N/A	N/A	N/A

Note: \* and \*\* denote statistical significance at the 5 and 1 per cent level respectively, p-values in square brackets

Table 6: EMU actual versus out-of-sample forecasts for Spanish three-month money market rates

	Average value 1991:1-2007:2	Projected Spanish to actual EMU rate
<i>Actual EMU three-month money market rate</i>	3.1	N/A
<i>Forecast for Spanish three--month money market rate</i>		
Linear model	6.0	2.0
Linear model with $rger_t$	5.7	1.9
L-STEEM	5.5	1.9
<i>Adjusted Forecast for Spanish three-month money market rate</i>		
Linear model	6.0	2.0
Linear model with $rger_t$	6.8	2.3
L-STEEM	4.5	1.4