evaluating currency crises: the case of the european monetary system

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Abstract

In this paper we examine the nature of currency crises. We ascertain whether the currency crises of the European Monetary System (EMS) were based either on bad fundamentals, or on self-fulfilling market expectations driven by external uncertainty, or a combination of both. In particular, we extent previous work of Jeane and Masson (2000) regarding evaluation of currency crisis. To this end we contribute to the existing literature proposing the use of three different Markov regime-switching models. Our empirical results suggest that the currency crises of the EMS were not due only to market expectations driven by external uncertainty, or `sunspots', but also to fundamental variables that help to explain the behaviour of market expectations.

Keywords: Currency crises, multiple equilibria, Markov-switching.

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1 Introduction

The currency crises of the EMS in 1992-1993, of Mexico in 1994 and the Asian crises in 1997 have been accompanied by considerable controversy over their causes. There are two main theoretical models that explain currency crises, the old currency crises model defined as seignorage-driven and the new generation currency crises model determined by self-fulfilling expectations of speculators. Advocates of the first generation model include Krugman (1979, 1996) and Flood and Garber (1984). Alternatively, advocates of the self-fulfilling view include Obstfeld (1996), Obstfeld and Rogoff (1995) and Cole and Kehoe (1995).

The logic of self-fulfilling crises is based on the idea that devaluation expectation increases the cost of retaining a peg and therefore the desire of the policy-maker to devalue. The most obvious way to do so is by increasing the ex-ante nominal interest rate, which affects economic growth negatively. Under such circumstances, the policy maker prefers to devalue rather than to maintain high interest rates even though she would have kept the exchange rate fixed if interest rates had been low. Therefore, the decision to devalue or not depends on market expectations regarding policy shifts in monetary policy pursued by a central bank. Jeane and Masson (2000) show that strategic complementarity between market expectations about the intended policy rule and the policy actually adopted produces multiple equilibria[1]. In particular, speculators' expectations about a devaluation force the policy-maker to revise the critical threshold of fundamentals that trigger devaluation[2]. Therefore, there are different levels of fundamentals where currency attack is imminent. Jeane and Masson (op. cit.) and Jeane (1997) assume that transition across different equilibria follows a Markov process independent of fundamentals and driven by extrinsic uncertainty. This is the case of sunspot equilibrium where devaluation expectation is the sum of the probabilities of devaluation in the next period weighted by the transition probabilities that the state of economy switch from the current to the

[1]Cooper and John (1988) shows that spillover and strategic complementarities give rise to multiple equilibria, the dynamics of which can be approached by regime switching.

[2]A vital assumption in this model is that speculators are rational and share a common knowledge of the information set. This is so, because, as Morris and Shin (1998) show, the absence of common knowledge gives rise to a unique equilibrium.

future states.

In this paper we examine the argument of Jeane and Masson (2000) that the transition probabilities across regimes are driven by extrinsic uncertainty, independent of macroeconomic variables. We argue that multiple equilibria do not only re°ect the reaction of monetary authorities to market expectations, but also re°ect their preferences regarding fundamental variables that might a®ect markets' expectations. More concretely, as Jovanovic (1989) shows if a sunspot is independent of fundamentals then it is necessary to distinguish the dynamic of fundamentals process from the sunspot process. A Markov regime-switching (MRS) model provides a framework that satis¯es the distinction between the two processes. In particular, the data generating process of a MRS model consists of two components: the ¯rst component gives rise to the autoregressive dynamic of fundamentals, and the second component describes the dynamics of an unobserved state variable which follows a Markov process. The latter component represent the extrinsic uncertainty (i.e. sunspot) which co-ordinates the public's expectations and leads the economy across di®erent equilibria. We argue that if the transition probabilities among equilibria are functions of fundamental variables (Filardo 1994; Diebold et. al. 1994), then sunspots are not independent of these fundamentals. Therefore, devaluation expectation might be driven at least to some extent by the fundamentals that a®ect the policy objectives of central banks.

In what follows, we utilise three di®erent MRS models to make an assessment of whether currency crises in the EMS were the result of deteriorating fundamentals or of self-ful¯lling expectations or a combination of the two. More concretely, we ¯rst use a MRS model with Autoregressive Conditional Heteroscedasticity (SWARCH) of the interest rate di®erential between the six individual member countries of the EMS and Germany. Secondly, we use a MRS model with dynamic time-varying transition probabilities. The reason for utilising a time-varying transition probability MRS model is to test whether the transition probabilities are driven endogenously by fundamentals. However, in a dynamic time-varying transition probability model we cannot estimate the fundamentals. Therefore, in the third speci¯cation we use a bivariate vector autoregressive MRS (BVAR-MRS) model to investigate possible variables that might a®ect the preferences of the central bank.

The paper proceeds by ¯rst giving a brief description of the theoretical underpinnings of currency crises. The empirical methodology adopted is explored in the subsequent section, with empirical ¯ndings reported and discussed in Section 4. The ¯nal section summarises and concludes.

2 Theoretical underpinning

The theoretical approach to credibility is based on the time inconsistency problem. A central bank minimises a loss function, which consists of two components: a) stabilisation of output-gap variability and b) stabilisation of in°ation variability. A policy of no in°ation is best in the long-run, but in the short-run the central bank has an incentive to in°ate. If the public expect zero in°ation, then the central bank faces the temptation to in°ate. The public knows about this incentive and adjusts its expectations accordingly to a positive in°ation rate. The result will be a positive in°ation rate without output expansion, since in°ation is completely anticipated. This in°ation bias is due to two reasons. First, the central bank has an incentive to in°ate; and second if the private agents expect zero in°ation then the marginal bene¯t of zero in°ation exceeds its marginal cost.

A solution to this in°ation bias requires a rise of the marginal cost of in°ation, as perceived by the central bank, to equal its marginal bene¯t. The ¯rst class of solution incorporates a reputation cost in a repeated-game version of the basic Barro-Gordon (1983a and 1983b) models. The second class of solution takes the form of institutional reforms that a government can adopt to lower in°ation expectations. One way to do so is to create an independent central bank that puts a high weight on in°ation stabilisation (see Rogo®, 1985). Another way is to appoint a central banker whose compensation is constructed so as to raise the marginal cost of in°ation (see Walsh, 1995). Finally, a third class of solution involves targeting rules that limit the ability of the central bank to stabilise demand and supply shocks. In particular, in a targeting rule framework a central bank is judged by its performance to achieve a prespeci¯ed level of some macroeconomic variables (i.e. in°ation)[3]. A ¯xed or a target-zone exchange rate system can be seen as

[3] For a recent discussion about in°ation targeting see Bernanke et. al. (1999), and Leidernman and Svensson (1995).

a targeting rule where the temptation to inflate is limited by the need to maintain an exchange rate target. More precisely, when lack of credibility is a problem for a central bank, pegging the exchange rate against the currency of a low inflation country helps to import credibility. Giavazzi and Pagano (1988) argue that by "tying their hands" the authorities of high inflation country can lower the output cost of disinflation. In view of this, the EMS was used as a mechanism to transfer credibility from Germany to other EMS countries.

In a fixed exchange rate system the loss function of the central bank incorporates an extra component that represents the cost of devaluation. In particular, a central bank minimises the following loss function:

$$L = \frac{1}{2}\lambda(y - y^n - k)^2 + \frac{1}{2}(\pi - \pi^*)^2 + C(\pi_t) \qquad (1)$$

where y = output, y^n = natural rate of output, π = inflation rate, π^* = inflation target which is assumed zero for simplicity, λ stands for the relative weight that the central bank attaches to output, relative to inflation stabilisation, and k is a parameter which may be justified on several accounts. Krugman (1996) modified the loss function above in terms of expected depreciation and deviation of the optimal value of exchange rate from its current level. We follow Krugman's representation and write (1) as:

$$L = [a(s^* - s^F) + b\varepsilon]^2 + R(\triangle\varepsilon) \qquad (2)$$

where s is the log of the exchange rate, s^* is the value that government would choose if it not receive any credibility cost, s^F is the parity to which it has stacked its reputation, ε is the expected rate of depreciation and R is a stochastic dummy variable that takes on value zero if the central bank does not devaluate and takes on value C if it does.

Jeane and Masson (2000) argue that the strategic complementarities between policymakers and speculators produce multiple fundamental based equilibria in which the economy moves across states with different levels of devaluation expectation. Jeane and Masson (op. cit.) also argue that devaluation expectations are independent of fundamentals and are driven by an external uncertainty (i.e. sunspot). In a sunspot equilibrium the economy can be in different states, where every state is characterised by different thresholds of fundamentals that currency crisis is imminent. Transition across states is governed by an unobserved state variable that follows a Markov process. Jeane and Masson (op. cit) assume that the transition probability matrix is constant across time (i.e. duration independent). Under such circumstances, shifts of the critical level of fundamentals reflect changes of market expectations and not changes of policy-maker preferences. Moreover, Jeane and Masson (op. cit) argue that the critical thresholds of fundamentals are not the same as those defined by the fundamental based equilibria models. The latter do not take into account market expectations about future state jumps which play an essential role in sunspot equilibria models. Therefore, in a sunspot equilibria, market expectation about future state shifts is an essential component of devaluation expectations.

In this study we investigate the assumption that the transition probability across the different states of the economy under consideration is independent of fundamentals. First, we use the Hansen (1992 and 1996) test for the number of states. Evidence of more than one state is consistent with the view that strategic complementarities between market agents and monetary authorities produce multiple equilibria, the dynamics of which can be approached by MRS models. Secondly, we utilise three different MRS models to evaluate the nature of currency crises. More concretely, we use a SWARCH model to estimate the ex ante probabilities of realignments on the basis of filtered probabilities. Moreover, we use a time-varying transition probability model and BVAR-MRS models, to investigate the impact of economic variables on the policy instruments of six EMS countries.

3 Econometric Methodology

3.1 A Univariate MRS Model

In what follows we use an econometric model that takes into account empirical regularities and theoretical considerations that suggests that expectation of devaluation goes through different regimes.[4] We employ a two-state SWARCH model of interest rate differential

[4] A number of studies (see Dahlquist and Gray, 2000; and Gray, 1996) have investigated the currency crises of the EMS in 1992 and 1993. The important conclusion of these studies is that devaluation expectation goes through different regimes (i.e. high and low credibility regime). However, none of these

between that of six individual EMS countries and of Germany. A SWARCH model of the interest rate differential can be expressed as:

$$i^D - i^G \mid I_{t-1} \sim \left\{ \begin{array}{l} N\left[b_0 + b_1\left(i^D - i^G\right)_{t-1}; h_{1t}\right] \\ N\left[c_0 + c_1\left(i^D - i^G\right)_{t-1}; h_{2t}\right] \end{array} \right\} \tag{3}$$

where i^D is the domestic EMS member country interest rate, i^G is the German interest rate, h_{st} (with $s = 1, 2$) is the variability of the error term u_t, defined as:

$$\begin{aligned} h_{1t} &= \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 \\ h_{2t} &= d_0 + d_1 u_{t-1}^2 + d_2 u_{t-2}^2 \end{aligned} \tag{4}$$

The subscripts 1 and 2 denote high and low credibility regime respectively. The transition between regimes is characterised by a (2×2) transition probability matrix $p = [p]_{ij}$, with $i, j = 1, 2$; where $p_{ij} = P(s_{t+1} = j \mid s_t = i)$. Moreover, every column of p sums to unity. This implies that $p'1 = 1$, where 1 is a (2×1) vector with unity elements. Equations (3) and (4) indicate that both the mean and variance of $(i^D - i^G)$ are subject to regime switching. A switch in the variance of $(i^D - i^G)$ might be due to an increase of the exchange risk premium that forces the domestic central bank to deviate from the monetary policy pursued by Germany. This is so because an increase of the variability of the interest rate differential will lead to higher inflation expectations and higher nominal interest rates. A higher interest rate will affect economic growth negatively forcing monetary authorities to devalue. Therefore, the conditional variance of the interest rate differential of the high credibility regime (i.e. regime 1) is expected to be lower than the conditional variance of the low credibility regime (i.e. regime 2).

Equations (3) and (4) are based on the theoretical framework introduced by Froot and Rogoff (1991). In particular, we assume that the interest rate differential between two countries follows an $ARMA(p, q)$ process of the form:

$$i^D - i^F = \alpha + A(L)\left(i^D - i^F\right) + \Delta S^E + B(L)\, e_t \tag{5}$$

studies used a formal statistic to establish the presence of two or more than two states.

where i^D is the domestic interest rate and i^F is the interest rate of a foreign country, ® is a constant, A (L) and B (L) are the AR and M A lag polynomials of order p and q, ¢S^E is the expected exchange rate change and e_t is the foreign exchange risk premium. Froot and Rogo® (1991, p. 297) argue that credibility can be measured by the sum of exchange risk premium and expected exchange rate change. Froot and Rogo® (op. cit.) also suggest that the error term in (5) can be thought of as the expectation of devaluation proxy. Under the assumption that the exchange rate follows a random walk (i.e. ¢S^E = 0) , the error term re°ects the credibility of monetary policy regarding an exchange rate target. Since credibility and reputation are established gradually over time, the conditional variance of the error term can be assumed to vary over time, which justi¯es modelling interest rate di®erentials as an ARCH process. However, evidence of high persistence both in the mean and in the variance of the interest rate di®erential (i.e. the sum of the coe±cient both of A (L) and B (L) is close to one) implies that there is a switch both in the mean and variance. The Hansen test and various model selection criteria were employed in this study to test for the number of states for each individual country. In all the cases the test strongly supports the presence of two states (see Section 4).

The evaluation of the character of currency crises under this speci¯cation draws on the ¯ltered probabilities of expected realignment. If the probability of the high credibility state declines before the time of a currency crisis this indicates that the currency crisis was the result of predictable deterioration of fundamentals. It was not self-ful¯lling. If the probability of a high credibility state decreases exactly at the time of the currency crisis then this is an indication of a self-ful¯lling currency crisis. This depends on whether the country under attack will adopt an in°ationary policy after the attack. In particular, if monetary authorities follow a tight monetary policy after the speculative attack (i.e. the probability of a high credibility regime increases after the crisis), then the currency crisis might not be due to self-ful¯lling expectations. Alternatively, if monetary authorities follow an in°ationary policy after the currency attack (i.e. the probability of a high credibility state remains low after the crisis), then the currency crisis is the result of self-ful¯lling expectations.

3.2 The Time Varying Transition Probability Model

The assumption of constant transition probabilities may be restrictive. Recently, Filardo (1994) and Diebold et. al. (1994) considered time-varying transition probabilities that were function of economic indicators. The implication of this specification is that fundamentals can help to predict future behaviour of the unobserved state variable. If the vector of economic fundamentals that determine the transition probabilities at time t is Z_t then the time-varying transition probabilities may have the following form:

$$p_{ij:t} = \exp\left(\beta_{ij;0} + Z'_{t-1}\beta_{ij:t}\right) \Big/ \left[1 + \exp\left(\beta_{ij;0} + Z'_{t-1}\beta_{ij:t}\right)\right]; \qquad i; j = 1; 2 \qquad (6)$$

Filardo (1998) shows that when modelling time-varying transition probabilities the information variables that affect time-variation must be conditionally uncorrelated with the unobserved state vector. In such a case the conditional maximum likelihood estimation should not be the same as the unconditional one and the information variables Z_t should be modelled jointly with the dependent variable (i.e. the interest rate differential). However, Chourdakis and Tzavalis (2000) suggest a dynamic specification of the time varying transition probability which guarantees that Z_t is uncorrelated with the unobserved state variable. Under such a specification, the vector of information variables contains lagged values of the error term u_t (i.e: $Z_t = \{u_{t-1}; u_{t-2}; \ldots\}$) : Therefore in a dynamic specification the time-varying transition probability model is written as:

$$p_{ij:t} = \exp\left(\beta_{ij;0} + \beta_{ij;1}u_{t-1}\right) \Big/ \left[1 + \exp\left(\beta_{ij;0} + \beta_{ij;1}u_{t-1}\right)\right]; \qquad i; j = 1; 2 \qquad (7)$$

In this paper, we follow the suggestion of Chourdakis and Tzavalis (2000) and we utilise (7) to make an assessment of the impact that the information variables have on the transition probabilities.[5] The type of news contained in the lagged value of the error

[5] We have also estimated with the specification proposed by Chourdakis and Tzavalis (2000). That includes the lagged value of $p_{ij:t}$ in the information set Z_t (i.e. $p_{ij:t} = \exp\left(\beta_{ij;0} + \beta_{ij;1}u_{t-1} + \beta_{ij;2}p_{ij:t-1}\right) \big/ \left[1 + \exp\left(\beta_{ij;0} + \beta_{ij;1}u_{t-1}\beta_{ij;2}p_{ij:t-1}\right)\right]$). However, the results regarding the coefficients $\beta_{ij;0}$ and $\beta_{ij;1}$ are similar to those derived from (7). Moreover, in the majority of the cases (i.e. the cases of Austria, Belgium, France and Netherlands) the coefficient $\beta_{ij;2}$ of $p_{ij:t-1}$ was not significant and the likelihood value function was lower than the likelihood value function of (7). Only in the cases of Italy and Spain the coefficients of $p_{ij:t-1}$ were significant with a

term in (7) (i.e. u_{t-1}) can be interpreted according to the sign of the parameters $\beta_{ij;1}$. In particular, if $\beta_{ij;1}$ is positive then $\partial p_{ij}/\partial u_{t-1} > 0$ which means that an increase of u_{t-1} increases the probability to switch to another state. Alternatively if $\beta_{ij;1}$ is negative then $\partial p_{ij}/\partial u_{t-1} < 0$, which implies that as u_{t-1} increases the probability of remaining in the same state increases. Regarding our theoretical priors, if the interest rate differential is in the high credible state (i.e. regime 1) then movement of u_{t-1} that might reflect demand and supply shocks will increase the incentive of monetary authorities to stabilise these shocks. This will increase the probability of switching to the low credibility state (i.e. regime 2). Consequently, in the high credible state the coefficient $\beta_{ij;1}$ is expected to be positive. Alternatively, if the interest rate differential is in the low credible state (i.e. regime 2) then changes of u_{t-1} increase the probability of remaining in the low credible state. Therefore the coefficient $\beta_{ij;1}$ is expected to be negative in the low credible state.

3.3 The Vector Autoregressive MRS (VAR-MRS) model

Although the dynamic time-varying transition probability model indicates that transition probabilities are endogenous and functions of economic indicators, it is not possible to identify the indicators. We adopt a bivariate Markov regime-switching model (BVAR-MRS) where the information variables and the interest rate differentials are estimated jointly. Under this specification we avoid the problem of endogeneity between transition probabilities and information variables and we can also identify which are the endogenous variables that affect the instrument of monetary policy (i.e. the interest rate differential).

The estimation procedure of the BVAR-MRS model is an extension of the basic VAR model. More precisely, consider a $p-$th order vector autoregression for the vector y_t:

$$y_t = c_0 + A_1 y_{t-1} + \dots\dots A_p y_{t-p} + u_t \tag{8}$$

where $u_t \sim IID(0; \sum)$ and $y_0; \dots; y_{t-p}$ are fixed. The most general specification of a M-state BVAR(M)-MRS model can be presented as follows:

$$y_t = v(s_t) + A_1(s_t) y_{t-1} + A_2(s_t) y_{t-2} + \dots + A_p(s_t) y_{t-p} + \sum u(s_t) \tag{9}$$

higher likelihood function than the likelihood value of (7).

where $s_t = 1, 2, \ldots m$ is a $(m \times 1)$ vector of state variable, $u \sim NID(0, I_n)$, $\sum(s_t)^{1/2}$ is the square root of state dependent variance covariance matrix and $A_i(s_t)$ is a state dependent $(n \times n)$ matrix of the autoregressive coefficients at $i - $th lag[6]. In our case A is a (2×2) matrix, and the order of the VAR is one. Therefore, we can write a two state bivariate first order MRS as follows:

$$y_t = c_{01,i} + A_{11,i} y_{t-1} + \Sigma_{i=1}^{2}(u_{t,i}) \tag{10}$$

or

$$\begin{bmatrix} \left(i^D - i^G\right)_t \\ x_t \end{bmatrix} = \begin{bmatrix} \alpha_{11,i} & \alpha_{21,i} \\ \alpha_{12,i} & \alpha_{22,i} \end{bmatrix} \begin{bmatrix} \left(i^D - i^G\right)_{t-1} \\ x_{t-1} \end{bmatrix} + \begin{bmatrix} \Sigma_{11,i} & \Sigma_{21,i} \\ \Sigma_{12,i} & \Sigma_{22,i} \end{bmatrix} u_t \tag{11}$$

where $i = 1, 2$ indicates the current regime and x_t denotes the vector of information variables[7].

A key choice among the macroeconomic variables that affect credibility is the overvaluation or undervaluation of exchange rates, which focuses on international competitiveness or on the current account (see Jean and Masson, 2000; Sarantis and Piard, 2000; Knot et al., 1998; and Caramazza, 1993). Other variables (inflation, unemployment, budget deficit) that appear to have an impact on the loss function of monetary authorities have also been considered as capable of explaining currency crises in the EMS (see De Grauwe, 1994; Masson, 1995; Rose and Svensson, 1994; and Arestis and Mouratidis, 2002). However, all of these studies use a single equation model to evaluate the credibility of monetary policy and the nature of currency crises in the EMS. The study by Arestis and Mouratidis (2002) is an exception, where two different MRS-bivariate vector autoregressive (BVAR-MRS) models are used to evaluate the credibility of nine EMS member countries based on the effects of inflation variability and output gap variability on monetary policy of the countries under consideration. In this study we use only the real exchange rate to evaluate

[6] Krolzig (1997) describes some particular specifications of BVAR-MRS models where the autoregressive parameters, the mean or the intercept are state dependent and the error term is homoscedastic or heteroscedastic.

[7] In the case of regime dependent heteroscedasticity (i.e. $\Sigma_1 \neq \Sigma_2$), Krolzig (1997) gives a comprehensive description of the Maximum Likelihood Estimation (MLE) of (11) for m number of states.

the currency crises. The set of macroeconomic variables could of course be extended to include fiscal variables. However, in most developed economies there is no mechanism to link deficit to money creation, and seignorage in the 1990s was negligible in the EMS (see De Grauwe, 1997).

4 Data and Results

Monthly averages of money rates were used in the following analysis with rates defined as the short-term borrowing rates affected between financial institution over the period March 1979 to April 1998. The starting point of the sample coincides with the inception of the EMS. Our data set includes six EMS member countries: Austria, Belgium, France, Italy, Netherlands and Spain; a seventh country, Germany, is used as a benchmark case. Not all these were member of the EMS for the full period of estimation; Austria joined the EMS in January 1995, continuing its policy of pegging the Schilling tightly to the DM, and Italy left the EMS after the crisis of 1992.

4.1 The SWARCH Model

Table 1 presents the results from a Hansen test regarding the number of states. The null hypothesis of linearity against the alternative of a Markov regime switching cannot be tested directly using a standard likelihood ratio (LR) test.[8] Therefore, we apply Hansen's standardized likelihood ratio test. This procedure requires an evaluation of the likelihood function across a grid of different values for the transition probabilities and for each state-dependent parameter. The value of the standardized likelihood ratio statistics along with the associated p-values under the null hypothesis are reported[9] in Table 1. We also test

[8] This is due to the fact that standard regularity conditions for likelihood-based inference are violated under the null hypothesis of linearity, as the parameters of transition probabilities are unidentified and scores with respect to parameters of interest are identically zero. Under such circumstances the information matrix is singular. However, appropriate test procedures that overcome the former or both of these difficulties do exist (Hansen, 1992, 1996; Garcia, 1998).

[9] The range [0:50; 0:99] in steps of 0:05 (10 grid points) is used as a grid for the transition probabilities; for the autoregressive coefficient and innovations variance, we use the range [0:1; 0:9] and [0:01; 0:17], respectively, in steps of 0:1 and 0:01 (9 grid points). The P-value is calculated according to the method

for the presence of a third state (Table 1). Furthermore, we have considered selection procedures based on the ARMA representation that a Markov-switching autoregressive process admits as well as procedures based on complexity-penalized likelihood criteria[10] (Table 1). The results provide strong evidence in favour of a two-states regime-switching specification for four countries. Exceptions to this are the cases of Austria and Italy where the results are rather contradictory, with the TPM selection criteria supporting a two-states dimension model while the Hansen test cannot reject the null hypothesis of linearity. However, in the case of Italy the presence of only one state is a rather strange result. This is so because Italy entered the EMS with high interest rate differential with respect to Germany and left the EMS after the currency crisis in 1992. Therefore in the case of Italy we should expect the presence of at least two regimes. In the case of Austria the acceptance of linearity would be consistent with the success of Austrian monetary authorities in shadowing the German monetary policy. We will treat the Austrian interest rate differential as a two state SWARCH model on the basis that Austria jointed the EMS in 1995.

Insert Tables 1

Table 2 reports the results of the SWARCH model for the six EMS countries. The coefficients b_i and c_i (where $i = 0, 1$) denote the constant and autoreregressive coefficient of the mean in the high credibility regime and in the low credibility regime respectively. The conditional variance is described by the coefficients α_i $(i = 0, 1, 2)$ for the high credibility regime and d_i $(i = 0, 1, 2)$ for the low credibility regime. Table 2 shows that the constant, when it is included in the conditional mean, is not significant in either regime. An exception to this result is the case of Italy where the constant is significant

described in Hansen (1996), using 1,000 random draws from the relevant limiting Gaussian processes and bandwidth parameter $M = 0, 1, \ldots, 4$; see Hansen (1992) for details.

[10] Using Monte Carlo analysis, Psaradakis and Spagnolo (2003) show that selection procedures based on the so-called three-pattern method (TPM) and the (Akaike, 1974) AIC are generally successful in choosing the correct state dimension, provided that the sample size and parameter changes are not too small. The BIC (Rissanen, 1978; Schwarz, 1978), and the HQC Hannan–Quinn criterion have a tendency to underestimate the state dimension. The good overall performance of the TPM method, combined with very low computational costs, makes it the best choice for selecting a lower bound for the number of Markov regimes in switching models.

in the low credibility regime (i.e. regime 2). This implies that there is convergence of the interest rate differential towards zero. Table 2 indicates that the persistence of both regimes, as measured by the conditional probabilities $p_{11} = 1 - p_{12}$ (i.e. high credibility regime) and $p_{22} = 1 - p_{21}$ (i.e. low credibility regime), is high, especially for the regime 1 (i.e. the high credibility regime). Moreover, the autoregressive coefficient in the high credible state is higher than that of the low credibility regime ($b_1 > c_1$). This is consistent with the view of Friedman and Laibson (1989) who show that small to moderate shocks are more persistent than large shocks. In particular, they argue that in the low credible regime where inflation expectations are high, the central bank will show its intention to reduce this expectation by exercising a strong upward pressure on the interest rates. This action leads to a large but not persistent change of interest rates, thereby relieving the pressure arising from inflation expectations. Finally, the conditional variance of the low credibility regime is higher than that of the high credibility regime (i.e. $\alpha_0 < d_0$) but less persistence (i.e. $\alpha_i < d_i$ where $i = 1, 2$). This implies that individual shocks fade away quickly in the low credibility regime (i.e regime 2) and last longer in the high credibility regime (i.e. regime 1).

Figure 1 presents the filtered probabilities that the interest rate differential is in the high credibility regime at the current period. Figure 1 indicates that only in the case of Italy the probability of being in the low credible state decreases before the crisis of September 1992. For the rest of the currencies the decrease in the probability of being in the high credibility regime (i.e. state 1) is contemporaneous to the crisis. This implies that the component of joint credibility of the EMS, argued by Rose and Svensson (1994), is very important. A contemporaneous jump in the probability of a switching state with the actual realignment is consistent with the Rose-Svensson (op. cit.) view that the currency crisis of 1992 had not been anticipated by the financial markets. At the same time it is consistent with their findings that the ERM realignment expectations are weakly related to macroeconomic variables[11].

The high value of the probability of being in the high credibility regime before and after the crisis of 1992, implies that the interest rate differentials of EMS countries do not

[11] Of the variables they examined, only inflation differentials between Germany and the EMS countries appear to affect ERM realignment expectations in a systematic way.

support the idea that the large devaluations of 1992 were the result of a self-ful¯lling crisis. Thus, it is clear that the crisis of 1992 had nothing to do with a self-ful¯lling speculative attack but, as has been discussed in De Grauwe (1994), there was a widespread agreement that Spain and Italy experienced a higher in°ation rate than the EMS average during 1987-1992. During this period, without any realignment, tensions had been building up for these two countries in the form of a growing loss of competitiveness[12]. Moreover, Eichengreen (2000) shows that in Italy, Spain, Sweden and UK unit labour cost had risen relative to their ERM partners before the crisis in 1992. Therefore, a deterioration of competitiveness was an important factor in the speculative attacks on these countries.

In the crisis of 1993 and the stormy period of 1995 the causes must have been di®erent because all the countries hit by those crises had current account surpluses and their currencies were not overvalued. Speculators in the currency markets attacked currencies that appeared not to be systematically overvalued, for example the Belgian franc, French franc and Danish krone. According to Gros and Thygesen (1998) and Eichengreen and Wyplosz (1993), the crisis of 1993 was the result of market expectations about the future change in the policy stance of France. France was in a di®erent cyclical position from Germany and thus, would like to follow a di®erent policy. Consequently, the crisis of 1993 might be consistent with the analysis of a self-ful¯lling speculative attack. Figure 1 shows that the probability of high credible state for the French and Belgian interest rate di®erential decreased in the middle of 1993, indicated by a depreciation of 3-4 percent. However, evidence that the probability of high credible state (i.e. state 1) increased after the crisis of 1993 for both countries questions the argument above that currency attack in 1993 was due to self-ful¯lling expectations. In general we can argue that although the crises in 1992 was due to the deterioration of fundamentals, the crises in 1993 might be a combination of market expectations and bad fundamentals.

[12] De Grauwe (1997) had reported that the currencies of these two countries were overvalued in 1992 by between 25 and 30 per cent and the choice for both of the countries was either to de°ate their economies or to use a large realignment.

4.2 The Time-Varying Transition Probability Model

Table 3 presents estimates from the time-varying transition probability model. The coe±cients of conditional mean and conditional variance are similar to the coe±cients of the SWARCH models. Therefore, b_i and c_i denote the coe±cients of conditional mean in the high and in the low credible state respectively. Moreover, the conditional variance of the high and of the low credibility regime is described by the coe±cients $®_i$ and d_i (where $i = 0; 1; 2$) respectively. The coe±cients $¯_{12;0}$ and $¯_{12;1}$ denote the constant and the slope coe±cient in the transition probability from the high credibility regime (i.e. regime 1) to the low credibility regime (i.e. regime 2). Alternatively, the coe±cients $¯_{21;0}$ and $¯_{21;1}$ indicate the constant and slope coe±cient of the transition probability from the low credibility regime to the high credibility regime. The main features of these results are consistent with those of the SWARCH models. In particular, the variance of the low credible state is many times greater than the variance of the high credible state but less persistence. (i.e. $d_0 > ®_0$ and $\sum_{i=1}^{2} d_i < \sum_{i=1}^{2} ®_i$). Moreover the autoregressive coe±cient of the interest rate di®erential in the high credible state is greater than the autoregressive coe±cient in the low credible state (i.e. $b_1 > c_1$).

To see how the transition probabilities vary within our sample period, in Figures 2 and 3 we present the graphs of transition probabilities from the high credibility regime to the low credibility regime (i.e. $p_{12:t}$) and from the low credibility regime to the high credibility regime (i.e. $p_{21:t}$). The graphs show high variability of transition probabilities especially of $p_{21:t}$ ¡ the transition probability from the low credibility regime to the high credibility regime. Figures 2 and 3 also show that the transition probability from the low credibility regime to the high credibility regime is higher than the transition probability from the high credibility regime to the low credibility state (i.e. $p_{21:t} > p_{12:t}$). This is more clear in the cases of France, Italy and the Netherlands where the transition probability $p_{21:t}$ oscillates close to unity and the transition probability $p_{12:t}$ oscillates close to zero than in the cases of Austria and Belgium. In the cases of Austria and Belgium, although $p_{21:t}$ is higher than $p_{12:t}$, both of them are close to zero especially after the last period of turbulence of the EMS in 1995. Finally, in the case of Spain the transition probability $p_{12:t}$ is similar to $p_{21:t}$. In particular, they are volatile in the period before 1986 and they become stable but very close to zero afterwards.

The estimates of the time-varying transition probability models show that in all cases, the coefficient $\beta_{ij;1}$ of u_{t-1} in (10) is significant with the exception of Belgium. These findings are consistent with the view that there are feedback effects between market expectations and central bank actions initiated by cycles of fundamental variables. This implies that transition across states is not independent of fundamentals as the sunspot models of currency crises suggest (Jeane and Masson , 2000). This is so because exogenous shifts in the agents' beliefs (i.e. sunspots) represented by an unobserved stochastic state variable are a function of economic indicators. In particular, fundamentals can be used to forecast future values of the state variable (see Filardo and Gordon, 1998). Therefore, market expectations are not independent of fundamentals.

The second column of Table 3 indicates that in the case of Austria the coefficient of u_{t-1} in (10) is significant in both states (i.e. $\beta_{12;1}$ and $\beta_{21;1}$). However, in the high credibility regime the coefficient $\beta_{12;1} = -3.943$ has the wrong sign. This might be due to the high credibility of Austrian monetary policy in shadowing German monetary policy. In such circumstances high credibility of monetary policy regarding future disinflationary policies has stabilising effects on the current policy (see Clarida et al., 1999). In the case of Belgium (see the third column of Table 2) neither of the coefficients $\beta_{12;1}$ and $\beta_{21;1}$ are significant. This implies that the speculative attack in 1993 against the Belgian franc might be due to the self-fulfilling expectations of speculators driven by a sunspot. However, this interpretation can be challenged on the ground that monetary policy did not become expansionary after the crisis in 1993. In particular, the filtered probability of being in the high credibility state increases after the crises in 1993[13]. The fourth column of Table 3 shows that in the case of France, both coefficients $\beta_{12;1}$ and $\beta_{21;1}$ have the correct sign but only the former is significant. This implies that the speculative attack on the French franc in 1993 was not due only to self-fulfilling market expectations driven by a sunspot but also due to fundamentals (i.e. supply or demand shocks) that might affect market expectations. In Italy only the coefficient $\beta_{12;1} = -4.967$ (i.e. the coefficient from the high credibility regime to the low credibility regime) is significant but with the wrong

[13] The filtered probabilities derived from the time-varying transition probability model are available from the authors upon request.

sign. This implies that fundamentals might not affect market expectations. However, as we mentioned in the previous sub-section, the probability of being in the high credible state decreases before the currency crisis in 1992. Therefore, in the case of Italy the speculative attack in 1992 was driven by fundamentals. In the Netherlands only in the high credibility regime is the coefficient of u_{t-1} significant and with the correct sign (i.e $\beta_{12:1} = 7.200$). As in the case of Austria, the wrong sign of coefficient $\beta_{21;1}$ might be due to the high credibility of the Dutch monetary authorities in their success to shadow the monetary policy pursued by Germany. Finally, in the case of Spain (i.e. the last column of Table 3) the coefficient of u_{t-1} is significant but with the wrong sign in the high credibility regime (i.e. $\beta_{12:1} = -1.416$) and insignificant in the low credibility regime. The implication of such evidence is that in the case of Spain either the credibility of the fixed exchange rate was high, or that agents' expectations are independent of fundamentals. However, based on the literature (see Eichegreen, 2000; and De Grauwe, 1997) and on the results of the previous section, we argue that the speculative attack on the Spanish Peseta in 1992 was due to the deterioration of competitiveness. Moreover, Spain moved gradually from an exchange rate targeting regime to an inflation targeting regime accompanied by transparency (i.e. better communication to the public). They, thus, managed to increase flexibility without undermining credibility. We conclude that the wrong sign of u_{t-1} in the high credibility regime (i.e $\beta_{12;1}$) is due to the high credibility of monetary policy followed by the Spanish central bank especially after the crisis of 1992 and an adoption of inflation targeting regime in 1994.

Evidence that the coefficient of u_{t-1} is significant in the majority of the countries under consideration indicates that fundamentals might be important determinant of the EMS currency crises in 1992 and 1993. However, in many occasions the coefficient $\beta_{ij:1}$ has the wrong sign. This implies that fundamentals alone cannot explain the behaviour of market expectations. Therefore, external uncertainty (i.e. sunspots) might affect market expectations in the period before the currency attack.

4.3 The BVAR-MRS model

We use a general BVAR-MRS model where the autoregressive parameters and the variance-covariance matrix are state dependent. We use only one lag since a higher number of lags do not increase the value of the likelihood function, and the results obtained by such speci¯cation have the same implication as those obtained using only one lag.

The coe±cients in Table 4 need to be explained. The subscript after the dot indicates the regime; the ¯rst subscript before the commas indicates the equation and the second the regressor under consideration. The coe±cient $c_{i0;1}$ denotes the constant coe±cient in the $i = 1; 2$ equation in the ¯rst regime (i.e. the high credibility regime); the coe±cient $a_{ij;1}$ is the coe±cient of the $j = 1; 2$ regressor in the $i = 1; 2$ equation in the ¯rst regime. Similarly, $c_{i0;2}$ is the constant coe±cient in equation $i = 1; 2$ in the second regime (i.e. low credibility regime); and $a_{ij;2}$ the coe±cient of $j = 1; 2$ regressor in the $i = 1; 2$ equation in the second regime. The variance covariance matrix $§_{ij:1}$ denotes the covariance between $i = 1; 2$ and $j = 1; 2$ in the ¯rst regime and the variance covariance matrix $§_{ij:2}$ is the covariance matrix in the second regime.

The likelihood value of the BVAR-MRS models is higher than that of the SWARCH and dynamic transition probability models. This implies that evaluation of the currency crisis model in a system framework is superior to that of a single equation model. The model is a reduced form and a structural interpretation of the result cannot be given[14]. We focus on the signi¯cance of the real exchange rate in the equation of the interest rate di®erential. The results presented in Table 4 are consistent with those of the SWARCH and time-varying transition probability models. First, the variance of the interest rate di®erential in the high credibility regime is lower than that of the low credibility regime [i.e. $§_{11;1} < §_{11;2}$]. Second, the autoregressive coe±cient of interest rate di®erential in the high credibility regime is higher than the coe±cient in the low credibility regime [$a_{11;1} > a_{11;2}$].[15] However, this is not true for the equation of the real exchange rate. For example, in the cases of Belgium and the Netherlands, the autoregressive coe±cient of the real exchange rate is the same in both regimes [$a_{22;1} \simeq a_{22;2}$]. Moreover, in the

[14] Although recently Owyang (2002) identi¯es VAR-MRS model imposing restrictions on the impulse response functions.

[15] In the case of Austria $®_{11;2} > ®_{11;1}$:

case of France the autoregressive coe±cient in the low credibility regime is higher than the autoregressive coe±cient in the high credibility regime [$a_{22;1} < a_{22;2}$]. These results indicate that in the case of France the regime generating process of the real exchange rate might be independent of the regime generating process of the interest rate di®erential[16]. This is also consistent with the evidence that in these two countries the variance of the real exchange rate in the low credibility regime is lower than the variance in the high credibility regime [$\S_{22;2} < \S_{22;1}$]. Third, in the majority of cases, in the equation of interest rate di®erential the coe±cient of real exchange rate is higher in the low credibility regime than that in the high credibility regime ($a_{12;2} > a_{12;1}$). This is in line with the view that the higher is the variability of real exchange rate the longer the economy will remain in the low credible state.

Table 4 indicates that in the cases of the Austria, Belgium and Spain (i.e. second, third and last column of Table 4), the real exchange rate is signi¯cant in the low credible state ($a_{12;2}$). This might be due to a number of reasons. For example, high in°ation expectations in this regime might increase the volatility of the real exchange rate, thereby a®ecting negatively the growth and competitiveness of the country under consideration. In the case of Netherlands (i.e. the sixth column of Table 4) the real exchange rate is signi¯cant in the high credible state ($a_{12;1}$). This implies that given a ¯xed nominal exchange rate in a free-shock environment (i.e. the high credibility regime), monetary policy is adjusted in line with the policy pursued by Germany.

In the cases of France and Italy (i.e. the fourth and ¯fth column of Table 4 respectively) the real exchange rate is not signi¯cant in any regime. However, the interest rate di®erential is signi¯cant in the equation of the real exchange rate. In the case of France the impact of the interest rate di®erential on the real exchange rate is signi¯cant in the low credible state (i.e. $a_{21;2}$). Evidence, as we mentioned above, that the variance of the real exchange rate in the low credibility regime is lower than that in the high credibility regime implies that the regime generating process of the real exchange rate is independent from the regime generating process of the interest rate di®erential. Under such circumstances we can test for Granger causality and ascertain the variable that drives the other

[16]Under such circumstances the transition probability matrix of the BVAR-MRS model is the Kronecker product of the transition probability matrix of each variable in the system (Philips, 1991).

from one regime to another (see Warne, 2000; Ravn and Sola, 1994; Sola, Spagnolo F. and Spagnolo N., 2002). Although we leave these tests for future work we are justifying in arguing that the interest rate differential might cause the real exchange rate. This is consistent with the competitiveness through disinflation policy followed by the French monetary authorities from 1983 to the early 1990s. That policy, however, failed in terms of its employment target (see Blanchard and Muet, 1993). Unemployment in France in the early 1990s was higher than the unemployment level before 1983.

In the case of Italy the interest rate differential in the equation of the real exchange rate is significant in the high credibility regime (i.e. $a_{21;1}$). Although, it is difficult to test for Granger causality between the interest rate differential and the real exchange rate since there is no evidence that the regime generating processes of the two variables are independent, we will argue that the former causes the latter to improve competitiveness. This is consistent with Eichengreen (2000) who shows that in Italy, unit labour cost relative to the country's ERM partners rose by seven percent between the beginning of EMS and the onset of the currency crisis in 1992.

Results from the BVAR-MRS models imply that issues like competitiveness, employment and growth are taken into account by the monetary authorities of Austria, Belgium, Spain and France. However, evidence that the filtered probability of being in the high credibility regime decreases at the time of the currency crisis, implies that it is difficult to predict the timing of the crisis.[17] Therefore, market expectations about the devaluation rule intended by the central bank, play a significant role during the period preceding the currency crisis.

5 Conclusions

In this paper we have used the MRS framework to analyse the nature of currency crisis. In particular, we extent previous work of Jeane and Masson (2000) regarding evaluation of currency crisis proposing the use of three different Markov regime-switching models. In particular, we have tested whether the currency crises of the EMS system were based on

[17] The figures of the filtered probability being in the high credible state derived from BVAR-MRS models are available from the authors upon request.

self-fulfilling expectations driven by external uncertainty or were the by-product of bad fundamentals or by the combination of both. We have covered the entire EMS in the cases of Austria, Belgium, France, Italy, the Netherlands and Spain. A seventh country, Germany, is used as a benchmark.

The evidence produced in this study shows that currency crises in the EMS were not based only on markets expectation regarding the threshold of fundamentals that force policy makers to devalue but also on the behaviour of these fundamentals.[18] Results from a SWARCH specification show that there is evidence of self-fulfilling speculation attack only in the case of Belgium and France in 1993. However, the fact that the filtered probability of a high credible state increases immediately after the currency attack, indicates that policy in these countries were very determined to maintain credibility. The time-varying transition probability model shows that fundamental variables have forecasting power regarding future states of the unobserved state variable. This implies that expectations were not driven by external uncertainty (i.e. sunspot). However, evidence that in some occasions the coefficient of the information variables (i.e. $\beta_{ij:1}$) has a wrong sign implies that the unobserved state variable is driven at least to some extent by external uncertainty (i.e. sunspot). Therefore both fundamentals and sunspots affect the behaviour of market expectations. Results of BVAR-MRS model show that real exchange rate has a significant effect on the policy makers loss function. However, in the case of France there is evidence that the regime generating processes of interest rate differential and of real exchange rate are either independent or there is only unidirectional causality running from the interest rate differential to the real exchange rate. This is an indication that the currency attack in France in 1993 might be due at least to some extent to self-fulfilling expectations independent of fundamentals.[19] The overall conclusion of this paper is that currency crises in the EMS can not be explained only on the basis of deteriorating fundamentals or on the basis of self-fulfilling market expectations but on the combination of the two.

[18] Summary of the results from each model is presented in the notes of the relevant Table.

[19] This is true provided that other fundamentals used in an BVAR-MRS models instead of real exchange rate have weak effects on the interest rate differential. Moreover, the latter Granger causes the fundamentals variables under consideration.

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TABLE 1

Standardized LR test	Austria	Belgium	France	Italy	Netherlands	Spain
	Linearity versus two-states Markov switching model					
LR	1:704	3:424	5:529	3:144	4:759	5:522
M = 0	(0:7900)	(0:0010)	(0:0001)	(0:1900)	(0:0001)	(0:0001)
M = 1	(0:8100)	(0:0030)	(0:0001)	(0:2900)	(0:0002)	(0:0001)
M = 2	(0:8600)	(0:0060)	(0:0001)	(0:3900)	(0:0002)	(0:0001)
M = 3	(0:9200)	(0:0060)	(0:0100)	(0:4200)	(0:0003)	(0:0001)
M = 4	(0:9400)	(0:0140)	(0:0100)	(0:5001)	(0:0003)	(0:0001)
	Two-states versus three-states Markov switching model					
LR		0:451	0:812		1:475	0:873
M = 0		(0:7216)	(0:4687)		(0:3524)	(0:4342)
M = 1		(0:7219)	(0:4691)		(0:3534)	(0:4351)
M = 2		(0:7225)	(0:4701)		(0:3534)	(0:4352)
M = 3		(0:7226)	(0:4717)		(0:3575)	(0:4355)
M = 4		(0:7302)	(0:4732)		(0:3589)	(0:4367)
Model selection criteria to determine the number of states						
Model Selection Criteria						
TPM	2	2	2	2	2	2
AIC	2	2	2	2	2	2
BIC	1	2	2	1	2	2
HQ	1	2	2	1	2	2

Note: See Hansen (1996) for details of the LR test statistic, such as the de¯nition of M. P -values are in parentheses. See Psaradakis and Spagnolo (2003) for a detailed description of the Model Selection Criteria adopted.

Table 2

Parameters estimates and related statistics for regime-switching ARCH models

Par.	Austria	Belgium	France	Italy	Netherlands	Spain
b_0			0:019 [0:524]	0:085 [0:202]	¡ 0:007 [0:687]	
b_1	0:889 [0:000]	0:690 [0:000]	0:983 [0:000]	0:997 [0:000]	0:932 [0:000]	0:975 [0:000]
c_0			¡ 1:364 [0:085]	¡ 0:796 [0:001]	¡ 0:171 [0:171]	
c_1	0:737 [0:000]	0:908 [0:000]	0:580 [0:003]	0:950 [(0:000]	0:585 [0:000]	0:971 [0:000]
®$_0$	0:179 [0:000]	0:154 [0:000]	0:253 [0:000]	0:294 [0:000]	0:213 [0:000]	0:342 [0:000]
®$_1$	0:694 [0:023]	0:063 [0:339]	¡ 0:002 [0:853]	0:239 [0:022]	0:218 [0:205]	0:407 [0:012]
®$_2$	0:237 [0:287]	0:039 [0:076]		0:001 [0:952]		
d_0	0:899 [0:000]	1:034 [0:000]	2:171 [0:000]	0:492 [0:002]	0:859 [0:000]	2:198 [0:000]
d_1	¡ 0:183 [0:000]	¡ 0:032 [0:488]	0:022 [0:780]	0:148 [0:158]	¡ 0:213 [0:000]	0:046 [0:063]
d_2	¡ 0:080 [0:257]	0:139 [0:000]		0:550 [0:029]		
p_{12}	0:059 [0:053]	0:049 [0:070]	0:077 [0:015]	0:100 [0:034]	0:026 [0:088]	0:016 [0:210]
p_{21}	0:129 [0:077]	0:033 [0:048]	0:358 [0:001]	0:343 [0:009]	0:149 [0:064]	0:036 [0:102]
LogLik	¡ 99:831	¡ 212:22	¡ 141:35	¡ 146:33	¡ 42:994	¡ 292:13

Notes: 1) The coe±cient b_j and b_j where j = 0,1 denotes the mean autoregressive coe±cient in regime 1 respectively and regime 2 respectively (i.e. high and low credible regime). Coe±cients ai and di where i = 0,1,2 denote the ARCH autoregressive coe±cients in regime 1 and 2 respectively. p12 and p21 denote the transition probabilities from regime 1 to regime 2 and from regime 2 to regime 1.

P-values are reported in squared brackets.

2) Summary of Results: Currency crises in 1992 were fundamental driven. In 1993 there is evidence of speculative attack in the cases of France and Belgium.

Table 3

Time-varying transition probabilities

Par.	Austria	Belgium	France	Italy	Netherlands	Spain
b_0	−0.022 [0.671]	−0.032 [0.128]	0.043 [0.156]	0.080 [0.196]	−0.007 [0.670]	0.114 [0.190]
b_1	0.893 [0.000]	0.672 [0.000]	0.987 [0.000]	1.000 [0.000]	−0.917 [0.000]	0.957 [0.000]
c_0	0.044 [0.459]	−0.509 [0.000]	−1.594 [0.000]	−0.917 [0.000]	−0.065 [0.669]	0.755 [0.000]
c_1	0.702 [0.000]	0.723 [0.000]	0.089 [0.562]	0.936 [0.000]	0.757 [0.000]	0.923 [0.000]
α_0	0.193 [0.000]	0.143 [0.000]	0.237 [0.000]	0.307 [0.000]	0.215 [0.000]	0.360 [0.000]
α_1	0.511 [0.001]	0.790 [0.125]	0.277 [0.002]	0.264 [0.004]	0.090 [0.125]	0.156 [0.000]
α_2	0.141 [0.008]	0.030 [0.253]	−0.004 [0.000]	−0.001 [0.959]	0.106 [0.110]	0.023 [0.288]
d_0	1.037 [0.000]	0.935 [0.000]	1.402 [0.010]	0.515 [0.000]	1.016 [0.000]	1.903 [0.000]
d_1	−0.230 [0.000]	0.036 [0140]	0.361 [0.024]	0.153 [0.164]	−0.150 [0.419]	0.071 [0.10]
d_2	−0.096 [0.384]	0.115 [0.019]	−0.050 [0.510]	0.631 [0.033]	−0.134 [0.041]	0.081 [0.000]
$\beta_{12;0}$	−2.872 [0.000]	−8.121 [0.087]	−2.524 [0.000]	−2.585 [0.000]	−4.315 [0.000]	−3.083 [0.000]
$\beta_{12;1}$	−3.943 [0.060]	−21.19 [0.094]	1.393 [0.035]	−4.967 [0.015]	7.200 [0.002]	−1.416 [0.027]
$\beta_{21;0}$	−1.663 [0.007]	−3.544 [0.000]	1.633 [0.158]	−0.530 [0.629]	−0.688 [0.599]	−5.059 [0.016]
$\beta_{21;1}$	−2.739 [0.023]	−0.394 [0.520]	−7.377 [0.253]	−19.16 [0.311]	14.75 [0.188]	−1.595 [0.094]
LogLik	−95.652	−199.069	−125.144	−136.997	−39.626	−276.526

Notes:1) $\beta_{12;i}$.where i = 0,1 denotes the constant and slope coefficient in the transition probability from regime 1 to regime 2 $\beta_{21;i}$ denotes the constant and slope coefficient in the transition probability from regime 2 to regime 1. P-values are reported in squared brackets.

2) Summary of results: Information variables have significant effect on the transition probabilities. However, in many occasionsthese effects have a wrong sign. Therefore, currency crises were due both to fundamentals and expectation driven by sunspots.

Table 4

Parameters estimates and related statistics for MS(2)-BVAR regime-switching model

Par.	Austria	Belgium	France	Italy	Netherlands	Spain
$c_{10;1}$			¡ 0:066 [0:894]	¡ 4:640 [0:102]	¡ 0:472 [0:000]	0:156 [0:877]
®$_{11;1}$	0:761 [0:000]	0:808 [0:000]	0:986 [0:000]	1:007 [0:000]	0:787 [0:000]	0:954 [0:000]
®$_{12;1}$	0:020 [0:174]	¡ 0:036 [0:232]	0:116 [0:849]	0:771 [0:100]	3:348 [0:000]	¡ 0:007 [0:979]
$c_{10;2}$			¡ 10:13 [0:122]	10:08 [0:116]	0:390 [0:699]	¡ 1:559 [0:019]
®$_{11;2}$	¡ 1:259 [0:134]	0:548 [0:000]	0:455 [0:008]	0:842 [0:000]	0:721 [0:000]	0:797 [0:000]
®$_{12;2}$	1:288 [0:031]	¡ 0:423 [0:000]	10:81 [0:175]	¡ 1:514 [0:147]	¡ 4:357 [0:714]	0:928 [0:000]
$c_{20;1}$			0:068 [0:006]	0:052 [0:417]	0:004 [0:167]	0:044 [0:254]
®$_{21;1}$	¡ 0:0002 [0:651]	0:0003 [0:137]	0:0001 [0:220]	¡ 0:0005 [0:02]	0:001 [0:307]	¡ 0:001 [0:003]
®$_{22;1}$	0:999 [0:000]	1:000 [0:000]	0:916 [0:000]	0:991 [0:000]	0:967 [0:000]	0:988 [0:000]
$c_{20;2}$			¡ 0:012 [0:668]	0:448 [0:307]	0:004 [0:533]	0:819 [0:002]
®$_{21;2}$	¡ 0:037 [0:122]	0:0001 [0:956]	0:0007 [0:063]	¡ 0:002 [0:448]	¡ 0:0006 [0:82]	0:002 [0:061]
®$_{22;2}$	1:017 [0:000]	1:000 [0:000]	1:018 [0:000]	0:928 [0:000]	0:967 [0:000]	0:764 [0:000]
§$_{11;1}$	0:108 [0:000]	0:370 [0:000]	0:0603 [0:000]	0:129 [0:005]	0:043 [0:000]	0:156 [0:000]
§$_{12;1}$	0:0001 [0:856]	0:0004 [0:021]	¡ 0:0003 [0:88]	0:0005 [0:070]	0:0001 [0:064]	0:0001 [0:788]
§$_{22;1}$	0:0008 [0:000]	0:0001 [0:000]	0:0001 [0:001]	0:0001 [0:000]	0:0001 [0:000]	0:0001 [0:001]
§$_{11;2}$	11:44 [0:856]	2:616 [0:000]	4:594 [0:118]	1:166 [0:021]	0:470 [0:000]	4:594 [0:118]
§$_{12;2}$	0:629 [0:000]	0:003 [0:713]	0:0001 [0:899]	0:010 [0:464]	0:001 [0:374]	0:0001 [0:899]
§$_{22;2}$	0:034 [0:859]	0:0002 [0:000]	0:00004 [0:001]	0:001 [0:013]	0:0001 [0:003]	0:0001 [0:000]
p_{12}	0:010 [0:873]	0:058 [0:014]	0:072 [0:029]	0:068 [0:191]	0:055 [0:114]	0:023 [0:157]
p_{21}	0:654 [0:574]	0:217 [0:003]	0:361 [0:001]	0:324 [0:023]	0:235 [0:040]	0:050 [0:208]
LogLik	756:45	722:22	728:08	641:75	900:82	489:66

Notes: 1) P-values are reported in squared brackets.

2) Summary of results: In the case of France there is evidence of self-ful¯lling currency attack driven by sunspots.

Figure 1: Filter Probabilities of High Credible State

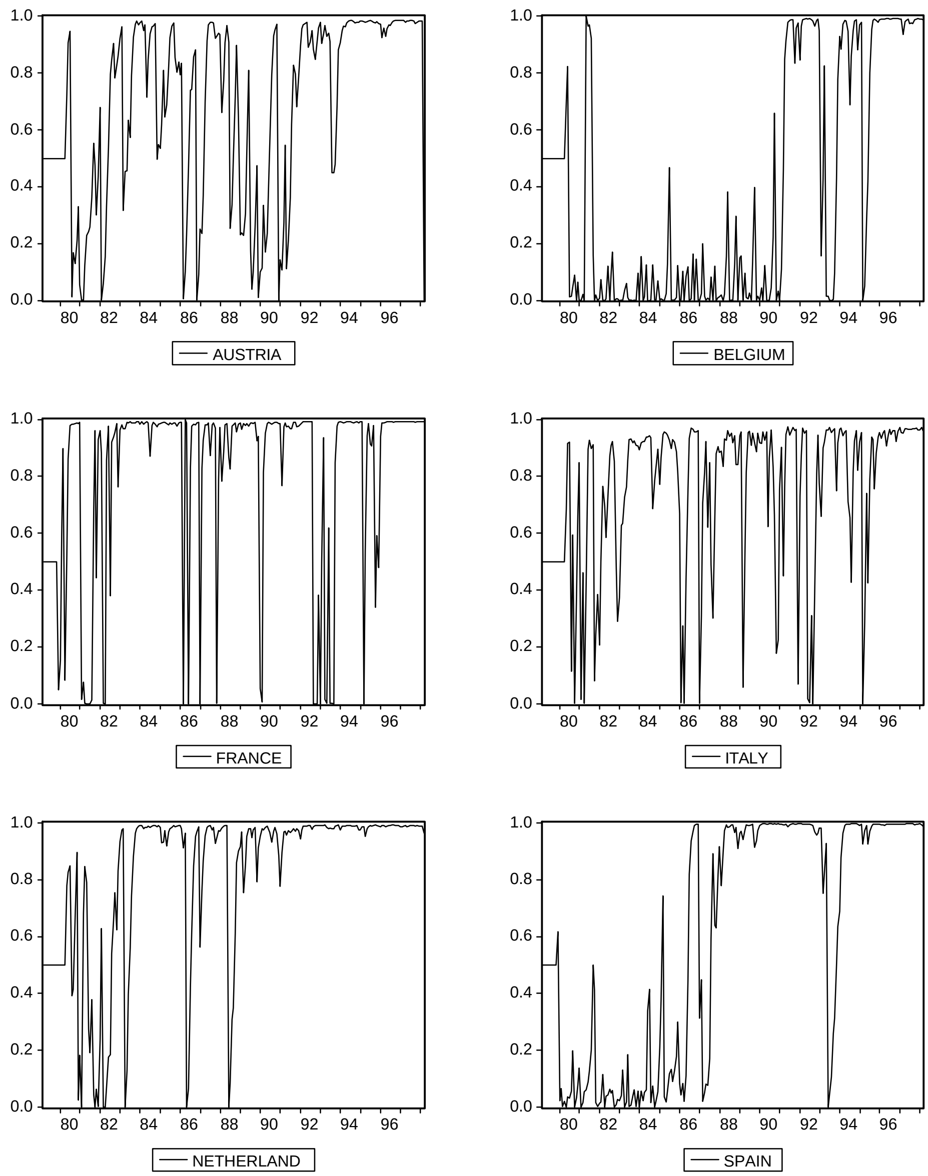
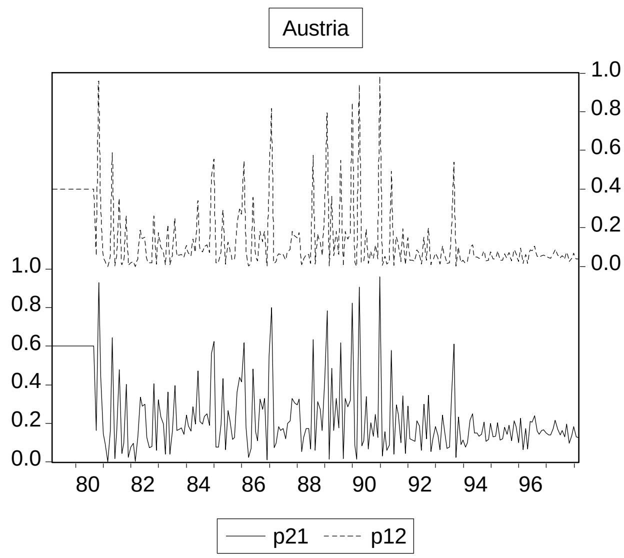
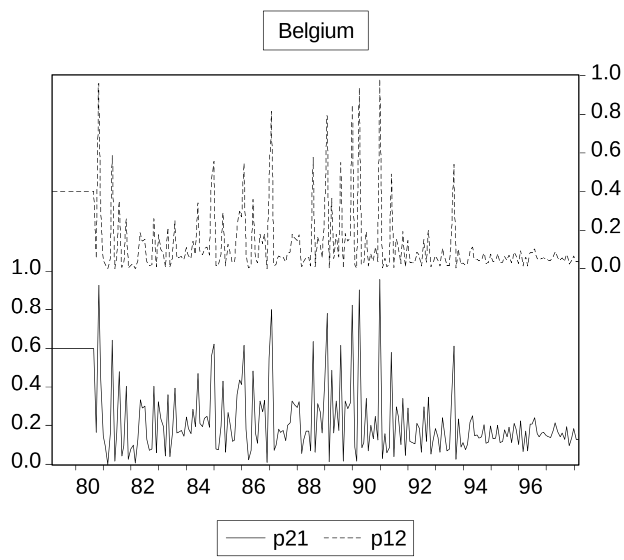


Figure 2: Time-varying Transition Probabilities





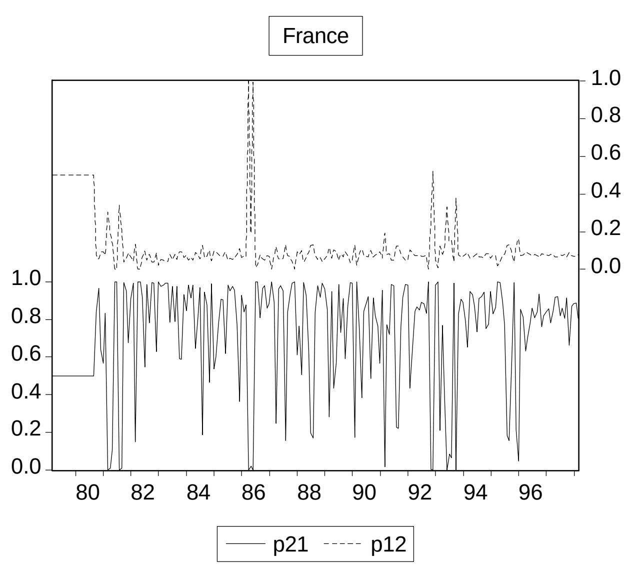


Figure 3: Time-varying Transition Probabilities