The Defense of Multilateral Exchange Rate Target Zones in the Face of Contagious Crises:

Lessons from the 1992 European experience

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SUMMARY

This paper investigate the way of defending a multilateral target zone system, such as the ERM–I, under the threat of a "contagious" speculative attack as the former may concern more than one member currency. Using the Kalman filter and applying the VAR approach, the Italian lira played a leading role during the months preceding the September 1992 crisis. This episode is very instructive since the resulting optimal defence of a multilateral exchange target zone appears in sharp constrast with very recent theoretical approaches. Furthermore, monetary authorities' interventions, although perfectly coordinated, may be highly unlikely, and thus fail given the unrealistic required amount, when attention is seriously paid to both their cost and the chosen horizon to achieve all the assigned exchange rate targets.

Keywords : Contagion, Currency Crisis, Target Zone, European Exchange Rate Mechanism

1. Introduction

The European Exchange Rate Mechanism (ERM) experienced one of its most turbulent period in 1992–1993. Although the system was originally designed to promote exchange rate stability among the participating currencies, this tranquillity broke into pieces in the mid-September 1992. The crisis erupted with the devaluation of the Italian Lira against all other currencies of the Mechanism on September 14th. It was immediately followed by an unprecedented speculative run on the British Pound and the Spanish Peseta during the "black Wednesday". On September 17th, both the Lira and the Sterling left the ERM, while the Spanish Peseta was devalued. Speculative pressure was then put on almost all the remaining currencies of the system, leading in particular to new devaluations of the Peseta, the Irish Punt, and the Portugese Escudo. One–way bets foreseen by investors only vanished with the widening of the fluctuation bands on August 2nd, 1993.

This singular episode of the ERM functioning has been already widely investigated. Indeed a growing number empirical studies aimed at identifying the effective roots of the September 1992 crisis. Although a special emphasis has been put to possible macroeconomic determinants of this sudden speculative attack, the empirical evidence is rather mixed[1].

This underlines the difficulties one faces at evaluating the credibility of an exchange rate fluctuation band. Econometric studies on the subject often rely on the literature on exchange rate target zones, renewed by Krugman's (1991) seminal article. However, empirical works based on the drift adjustment method, initially suggested by Rose and Svensson (1991), lead to unappealing conclusions[2]. Alternative approaches, such as Edin & Vredin's (1993) probit analysis, may give more convincing results despite their own methodological issues[3]. However almost all previous works concerned fluctuation bands of **bilateral** exchange rates defined against the German currency. What is surprising is that authors fail to take a striking feature of the ERM into account: its **multilateral** nature. Possible interactions among the set of bilateral fluctuation margins that represents the ERM were completely neglected until Eichengreen, Rose

[1] See the detailed analysis provided by Eichengreen and Wyplosz (1993). Chen and Giovannini (1993) and Rose and Svensson (1994) also give insights of the influence of macroeconomic developments on the credibility the ERM bilateral bands.

[2] In this spirit, Weber (1991) applied the Kalman filter principle to evaluate the credibility of ERM bands. More recently, Avouyi-Dovi and Laffargue (1994) have performed a Kalman filter estimation of parameters of the assumed Poisson process followed by the central parity.

[3] Unfortunately, their Probit model does not take into account the chronological order of the data.

& Wyplosz (1995) influential paper.

The issue raised by the aforementioned studies is even more acute given the vast literature related to the asymmetric functioning of fixed exchange rate systems[4]. In this view Vector Auto–Regressive (VAR) specifications have been used in order to search for interest rates linkages within the European Monetary System[5]. But such econometric models are rarely applied to groups of *more than two* bilateral exchange rates. This is even more surprising given the emphasis put by a number of theoretical analyses, like De Grauwe (1977), Aoki (1977) or Canzoneri (1982), on the destabilizing role of defending a particular bilateral parity on the others when at least three currencies are placed under a (common) fixed exchange rate regime.

It remains that serious attention has been paid to the risk of contagion of currency crises only after the South East Asian tumoil in 1997–98. . While these authors reached quite different conclusions about the transmission channels[6] during the European turmoil in 1992-93, their econometric methods fall under severe criticism[7].

The main purpose of this study is to give another look at the 1992 events with the aim at focusing on the conduct of concerted official interventions to defend a multilateral target zone system under the threat of a speculative run. To this end, the extent to which the September 1992 crisis was expected by market participants is first of all reassessed. Then it is shown how bilateral exchange rates within the ERM interacted in the few weeks before the devaluation of the Lira. Suggestive guidelines are finally fomulated as concerned the optimal intervention policy to defend a multilateral target zone system facing a contagious speculative attack. The simulated strategies based on our econometric model challenged the conclusions drawn from theoretical works pioneered by Flandreau (2000) and Serrat (2000).

In section 2, we develop a new methodology to assess the credibility of exchange rate bands. In the perfectly credible target zone model, it is well-known that the *expected* exchange rate cannot go beyond the fluctuations limits provided that expectations are formed rationally. The Kalman filter principle is used to infer the probability for a bilateral rate to lie outside its band at a given date. This methodology is then applied to the mark/lira (DEM/ITL) rate in its narrow band. It appears that tensions

[4]See, for example, de Grauwe (1989) and Fratianni and von Hagen (1990) for a theoretical treatment of asymmetries in the European Monetary System.

[5]See, for example, Katsimbris and Miller (1994).

[6]Drazen (1999), and Favero & Giavazzi (2000) stress very different mechanisms. See also Dornbusch, Park & Claessens (2000) for an overview on this growing subject.

[7]See Pericoli and Sbracia (2003) and Rigobon (2001, 2003) for an extensive discussion.

became appreciable since mid-August 1992. In the third section, we ask whether specific features of the DEM/ITL process contributed to the September 1992 exchange rate crisis. None of the considered individual sources help explain the DEM/ITL path. This last observation leads us to distinguish, in section 4, between specific and common factors which might have influenced intra-ERM exchange rates against the German mark from October 8th, 1990 (which coincides with the entry of the GBP into the ERM) to September 11th, 1992 (*i.e.* just before the ITL devaluation). Our Kalman filter estimations show evidence of common evolutions of ESP, ITL, and GBP against the mark only after the Danish no to the Maastricht Treaty. Following Lütkepohl (1991),we then use the VAR methodology to investigate for causal relationships among these three bilateral rates. There is evidence of a leading role for the DEM/ITL rate in the scenario of the September 1992 crisis. Finally, our time-invariant state-space representation proves to be well-suited for analyzing the behavior of these exchange rate returns in the steady-state. We then draw interesting conclusions on exchange rate policy targeting. In the final section 5, we draw conclusions and suggest some directions for future research.

2. Evaluating credibility problems under a target zone regime

In this section, we attempt to provide a reliable indicator of the bilateral DEM/ITL band credibility. Our procedure is divided in two steps. First, we get an estimate of the DEM/ITL exchange rate process when it evolved within its narrow fluctuation margins. We apply the Kalman filter methodology which appears particularly appropriate in the context of rational expectations, a traditional hypothesis in target zone model following Krugman's (1991) seminal work. Next, we suggest a new approach to extract a potentially useful measure of tensions in an bilateral exchange rate band. Applying this original methodology to the DEM/ITL experience shows that the September 1992 crisis may well have been anticipated by exchange rate market participants. This seems to contradict conclusions drawn from most of past empirical studies on this troubled period of the ERM.

2.1. Estimation of the exchange rate process.

Let s_t the logarithm of the exchange rate. We consider that s_t may be modelled as the following ARMA(2,2) process:

$$s_t = \mu + \phi_1 s_{t-1} + \phi_2 s_{t-2} + u_t - \theta_1 u_{t-1} - \theta_2 u_{t-2} \qquad (2.1)$$

with $u_t \sim \mathcal{N}(0, \sigma^2)$.
We assume here that DEM/ITL movements during the 1990-1992 period can be reasonably described by such a process[8].

2.1.1. Principle of the Kalman filter.

We use HARVEY's (1989) notations. A state-space model is defined by a transition equation and a measurement equation. In the former equation we postulate the relationship between an observable vector and a state vector, while the latter equation describes the generating process of state variables.
Here, we suppose that the state vector $\boldsymbol{\alpha}_t$ is generated by a first-order Markov process of the form:

$$\boldsymbol{\alpha}_t = T_t \boldsymbol{\alpha}_{t-1} + \mathbf{c}_t + R_t \boldsymbol{\eta}_t, \quad t = 1, \ldots, T \tag{2.2}$$

where $\boldsymbol{\alpha}_t$ is the m-dimension state vector, T_t is a $m \times m$ matrix, $\mathbf{c}_t$ is a $m \times 1$ vector, R_t is a $m \times g$ matrix.
The measurement equation of our state-space representation is:

$$\mathbf{y}_t = Z_t \boldsymbol{\alpha}_t + \mathbf{d}_t + \boldsymbol{\epsilon}_t, \quad t = 1, \ldots, T \tag{2.3}$$

where $\mathbf{y}_t$ is a univariate time series $N \times 1$, Z_t is a $N \times m$ matrix, $\mathbf{d}_t$ is a $N \times 1$ vector. $\boldsymbol{\eta}_t$ and $\boldsymbol{\epsilon}_t$ are assumed to be white processes of dimension $g \times 1$ and $N \times 1$ respectively. These two last uncorrelated processes are Gaussian with zero mean and with respective covariance matrix Q_t and H_t, that is:

$$\begin{array}{ccc} \mathrm{E}(\boldsymbol{\eta}_t) = 0 & \text{and} & \mathrm{Var}(\boldsymbol{\eta}_t) = Q_t \\ \mathrm{E}(\boldsymbol{\epsilon}_t) = 0 & \text{and} & \mathrm{Var}(\boldsymbol{\epsilon}_t) = H_t \end{array} \tag{2.4}$$

First of all, we need to initialize the Kalman filter. This can be done by assuming that the initial position is a Gaussian variable such that:

$$\mathrm{E}(\boldsymbol{\alpha}_0) = \mathbf{a}_0 \quad \text{and} \quad \mathrm{Var}(\boldsymbol{\alpha}_0) = P_0 \tag{2.5}$$

If we consider now $\mathbf{a}_t$ the optimal estimator of α_t based on all the relevant and available observation at time t, we have:

$$\mathbf{a}_t = \mathrm{E}_t[\boldsymbol{\alpha}_t] \tag{2.6}$$

[8] Although this specification of exchange rate movements appears rather arbitrary at first hand, it appears to be well-suited for describing daily DEM/ITL variations. Given the frequency of our observations, no macroeconomic variables (such as money stock or economic activity) can be included in this specification, in departure from most of theoretical models of exchange rate determination.

where E_t indicates the conditional expectation operator. The covariance P_t of this estimator is defined by:

$$P_t = \mathrm{E}_t \left[(\mathbf{a}_t - \boldsymbol{\alpha}_t) (\mathbf{a}_t - \boldsymbol{\alpha}_t)^\top \right] \tag{2.7}$$

The Kalman filter consists in the following set of recursive equations:

$$\begin{cases} \mathbf{a}_{t|t-1} = T_t \mathbf{a}_{t-1} + \mathbf{c}_t \\ P_{t|t-1} = T_t P_{t-1} T_t^\top + R_t Q_t R_t^\top \\ \widetilde{\mathbf{y}}_{t|t-1} = Z_t \mathbf{a}_{t|t-1} + \mathbf{d}_t \\ \mathbf{v}_t = \mathbf{y}_t - \widetilde{\mathbf{y}}_{t|t-1} \\ F_t = Z_t P_{t|t-1} Z_t^\top + H_t \\ \mathbf{a}_t = \mathbf{a}_{t|t-1} + P_{t|t-1} Z_t^\top F_t^{-1} \mathbf{v}_t \\ P_t = \left(I_m - P_{t|t-1} Z_t^\top F_t^{-1} Z_t \right) P_{t|t-1} \end{cases} \tag{2.8}$$

$\mathbf{a}_{t|t-1}$ and $P_{t|t-1}$ are the best estimators of $\boldsymbol{\alpha}_t$ and P_t (respectively), based on the information available at time $t-1$. $\mathbf{v}_t$ is the innovation process with covariance matrix F_t. I_m indicates the identity matrix of dimension m.
Let θ the vector of parameters. The log-likelihood function can be expressed in terms of the innovation process. It is then equal to:

$$\ell_t = \log L_t = -\frac{N}{2} \log 2\pi - \frac{1}{2} \log |F_t| - \frac{1}{2} \mathbf{v}_t^\top F_t^{-1} \mathbf{v}_t \tag{2.9}$$

Let $\ell(\mathbf{y}, \theta) = \sum_{t=1}^{T} \ell_t$. $\hat{\theta}_{\mathrm{ML}}$ is the maximum likelihood estimator for the sample $\mathbf{y}$ if we have:

$$\ell\left(\mathbf{y}, \hat{\theta}_{\mathrm{ML}}\right) \geq \ell(\mathbf{y}, \theta) \qquad \forall \theta \in \Theta \tag{2.10}$$

where Θ represents the parameters space.

Having briefly recalled the essence of Kalman filtering, we now turn on the state-space formulation of the DEM/ITL process.

2.1.2. State-space representation.

From what follows, the ARMA(2,2) process (2.1) can be put in the following state-space form:

$$\begin{cases} s_t = \begin{bmatrix} 1 & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} s_t \\ s_{t-1} \\ u_t \\ u_{t-1} \end{bmatrix} \\ \begin{bmatrix} s_t \\ s_{t-1} \\ u_t \\ u_{t-1} \end{bmatrix} = \begin{bmatrix} \phi_1 & \phi_2 & -\theta_1 & -\theta_2 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} s_{t-1} \\ s_{t-2} \\ u_{t-1} \\ u_{t-2} \end{bmatrix} + \begin{bmatrix} \mu \\ 0 \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} 1 \\ 0 \\ 1 \\ 0 \end{bmatrix} \eta_t \end{cases} \tag{2.11}$$

with $\eta_t \sim \mathcal{N}(0, \sigma^2)$. ϕ_1, ϕ_2, θ_1, θ_2, μ and $\sigma^2$are the coefficients to be estimated. We have performed a maximum likelihood estimation of the DEM/ITL exchange rate.

We use *daily* observations, available from Reuter services, from January 8th, 1990 to September 11th, 1992. Figure **2.1** shows DEM/ITL movements during this period.

Fig. 2.1 about here

Table **1** below put together the estimated values of these parameters and their corresponding significance.

Table 1

Parameter	Value	Std. error	t-Student	p-value
μ	-0.001049	0.0023750	-0.44	0.66
ϕ_1	1.719989	0.000353	4871.90	0.00
ϕ_2	-0.720504	0.000825	-873.36	0.00
θ_1	0.690886	0.033169	20.83	0.00
θ_2	0.051160	0.033077	1.55	0.12
σ	0.001196	0.000032	37.74	0.00

The DEM/ITL rate seems to have evolved as a pure ARMA(2,1) process. Note that p-values lead to the rejection of the statistical significance of μ and θ_2 at the 10 percent risk level. Parameters of the AR component have opposite signs. This induces conflicting influences of past exchange rate realizations on the current value of the relative

price of the mark against the lira.
We will pay more attention to this singular feature in the following paragraphs. Before this, we would like to investigate whether this state-space representation is stable over time or not.

2.1.3. Checking for parameter instability.

NYBLOM (1989) has built a econometric procedure to test the time-invariance of parameters over time. An interesting extension of this procedure is the one suggested by HANSEN (1990) who extends Nyblom's methodology to non-linear models. To this end, he develops Lagrange multiplier tests to evaluate the instability of the model coefficients. We implement both procedures to identify potential structural breaks in the sample considered.
Nyblom considers the statistic L defined by:

$$L = \operatorname{tr}\left[J^{-1}\sum_{j=1}^{T}\left(\sum_{t=j}^{T} d_t\right)\left(\sum_{t=j}^{T} d_t^{\top}\right)\right] \tag{2.12}$$

with ℓ_t' the first-order derivative of the log-likelihood function ℓ_t:

$$\ell_t' = \frac{\partial\,\ell_t}{\partial\,\theta} \tag{2.13}$$

Let $I(\theta)$ be the information matrix[9]. For the matrix J, we have chosen:

$$J = I(\theta) \tag{2.14}$$

The critical value of Nyblom's test $\mathbf{t}^N = \frac{L}{T^2}$ is tabulated on page 227 of NYBLOM (1989). Hansen has proposed a modified version of Nyblom's test in order to take non linearities into account. He constructs the variable L^* which is given by the following formula:

$$L^* = \sum_{t=1}^{T} S_t^{\top} I(\theta)^{-1} S_t \tag{2.15}$$

S_t is evaluated at the maximum of the log-likelihood function, that is:

$$S_t = \sum_{j=1}^{t} \ell_j'\Big|_{\theta=\hat{\theta}_{\mathrm{ML}}} \tag{2.16}$$

[9] To get an estimate of the information matrix, we have systematically computed the opposite of the Hessian of the log-likelihood function at the optimum.

Hansen proposes a test for the constancy of each parameter. It is based on:

$$L_k^* = \sum_{t=1}^{T} \frac{S_{k,t}^2}{I_{k,k}(\theta)} \quad \text{for all} \quad k = 1, ..., K \tag{2.17}$$

where K is the number of parameters to be estimated. Both statistics L^* and L_k^* follow the same probability law as L. Given this property, we can build the two statistics $\mathbf{t}^H = \frac{L^*}{T^2}$ and $\mathbf{t}_k^H = \frac{L_k^*}{T^2}$ which have the same critical values as $\mathbf{t}^N$.

Table 2

Parameter	$\mathbf{t}_k^H$
μ	0.249
ϕ_1	0.250
ϕ_2	0.250
θ_1	0.037
θ_2	0.043
σ_ε	0.182

The statistics $\mathbf{t}^N$ and $\mathbf{t}^H$ are equal to 1.2398 and 1.2415 respectively. The critical values are 1.686 at the 5% risk level and 2.117 at the 1% level. This implies that there is global time-invariance of the chosen representation during the whole period.
For the statistic $\mathbf{t}_k^H$, the critical values are 0.461 at the 5% level and 0.743 at the 1% one. *Results reported in table 1 above show a strong acceptance of the null hypothesis of constancy for each parameter.*
We would like to show evidence on specific properties of DEM/ITL fluctuations within the narrow band.

2.1.4. Evaluating some parameter restrictions.

We have performed a Wald test to check the validity of some restrictions in our econometric specification. This test is based on the likelihood principle. Consider the null hypothesis $\mathrm{H}_0 : R(\theta) = 0$ such that:

$$\begin{array}{rccc} R: & \Theta & \longrightarrow & \mathbb{R}^g \\ & \theta & \longmapsto & R(\theta) \end{array} \tag{2.18}$$

g represents the number of equality constraints to be satisfied. Let $I(\theta)$ be the information matrix, T be the number of observations and $\hat{\theta}_{\mathrm{MV}}$ be the estimator obtained by the maximum likelihood method. Under the null hypothesis, it is possible to establish that the Wald statistic $\mathbf{W}$ behaves asymptotically as a χ_g^2, that is:

$$\mathbf{W} = TR\left(\hat{\theta}_{\mathrm{MV}}\right)^{\top}\left[\frac{\partial R}{\partial \theta^T}\left(\hat{\theta}_{\mathrm{MV}}\right) I\left(\hat{\theta}_{\mathrm{MV}}\right)^{-1} \frac{\partial R^T}{\partial \theta}\left(\hat{\theta}_{\mathrm{MV}}\right)\right]^{-1} R\left(\hat{\theta}_{\mathrm{MV}}\right) \tag{2.19}$$

with

$$\mathbf{W} \underset{as}{\sim} \chi_g^2$$

Given our state-space specification (2.1), it is interesting to test the following equality constraint:

$$\phi_1 + \phi_2 = 1 \tag{2.20}$$

If we cannot reject the above assumption, it is possible to rewrite the state equation of our state-space model as it will be shown below.

The value of the Wald test is 0.1907 and its corresponding p-value is 0.6623. This implies that *it is not possible to reject the null (2.20) for the DEM/ITL rate during the global period*. Moreover, estimates of μ and θ_2 in table **1** are not significantly different from the zero value. From all these observations, the conditional expected rate of variation of the exchange rate $E_{t-1}\left[s_t - s_{t-1}\right]$ can be modelled as:

$$E_{t-1}\left[s_t - s_{t-1}\right] = -\phi_2\left[s_{t-1} - s_{t-2}\right] - \theta_1 u_{t-1} \tag{2.21}$$

If we omit the term u_{t-1}, a variation of the exchange rate at period $(t-1)$ is followed by a movement in this rate at the current period t in the same direction due to the negative sign of the ϕ_2 estimate. Thus, it looks as if the DEM/ITL rate were submitted to *a correction mechanism*, suggesting the following generating process for this exchange rate:

$$s_t = s_{t-1} + \rho_t\left(s_{t-1} - s_{t-2}\right) + \varepsilon_t \tag{2.22}$$

ρ_t represents the degree with which past fluctuations are counterbalanced (provided that ρ_t has a negative sign). This specific feature of DEM/ITL movements under the standard ERM regime will be investigated in further details in the next section.

What appears to be difficult is to show the turbulences which perturbed the price of lira relative to the German currency just before the September 1992 parity adjustment. In the next paragraph, we suggest an alternative way of inferring the strength with which these tensions exercised on the DEM/ITL rate.

2.2. How to evaluate the strength of tensions in the ERM?

As it has already been signalled in the introduction, one crucial problem, which previous empirical works on target zones credibility had to deal with, is to give a consistent evaluation of the realignment risk inside a zone. In a first time, the uncovered interest parity (UIP) relationship which prevails in standard target zone models, has been exploited to get an estimate of the expected rate of realignment[10]. Other authors have suggested an alternative two-step procedure which is based on the following principle:

1. first, get an estimate of the realignment probability,
2. then evaluate the expected size of the central parity shift.

Given these two measures, one can infer the expected rate of realignment[11]. Others finally tried to link the expected rate of a jump in the central parity to macroeconomic factors. Although interesting in its spirit, this last procedure gives poor results. Besides, it hardly explains the September 1992 events in the ERM[12].

Here, we would like to provide an alternative indicator of the credibility of an exchange rate band. Given KRUGMAN's (1991) theoretical treatment, it appears that under the rational expectation hypothesis, coupled with the UIP relation, the perfect credibility of the band implies that the **expected** exchange rate cannot lie outside its prescribed fluctuation margins. This is why we try to evaluate the probability for the DEM/ITL rate to be expected to lie outside its limits *at a given date*. One of the merits of this approach is that we do not need to assume that the UIP is valid, contrary to WEBER's (1991) and MUNDACA's (1995) regime-switching approaches[13]. The Kalman filter is particularly appropriate to tackle such a problem.

An interesting feature of state-space models is that they can be simulated. To run simulations, we consider $\tilde{\boldsymbol{\eta}}_t$ and $\tilde{\boldsymbol{\varepsilon}}_t$ the respective realizations of the two noise variables $\boldsymbol{\eta}_t$ and $\boldsymbol{\varepsilon}_t$. We generate sequences of realizations of the process $\mathbf{y}_t$ by using the following pair of recursive equations:

$$\begin{cases} \tilde{\boldsymbol{\alpha}}_t = T_t \tilde{\boldsymbol{\alpha}}_{t-1} + \mathbf{c}_t + R_t \tilde{\boldsymbol{\eta}}_t \\ \tilde{\mathbf{y}}_t = Z_t \tilde{\boldsymbol{\alpha}}_t + \mathbf{d}_t + \tilde{\boldsymbol{\varepsilon}}_t \end{cases} \tag{2.23}$$

[10] See the pionering studies of SVENSSON (1993) as well as ROSE and SVENSSON (1991) on this subject.

[11] EDIN and VREDIN (1993) have built a Probit model to evaluate the reaction of monetary authorities.

[12] See CHEN and GIOVANNINI (1993) and ROSE and SVENSSON (1994).

[13] Relying on the UIP assumption, AVOUYI-DOVI and LAFFARGUE (1994) have performed a Kalman filter estimation of parameters of the Poisson process followed by the central parity.

with $\tilde{\boldsymbol{\alpha}}_0 = \mathbf{a}_0$.

We can use a Cholesky decomposition to simulate the stochastic terms. Suppose that Σ is the covariance matrix and μ the mean vector of the couple $(\boldsymbol{\eta}_t,\boldsymbol{\varepsilon}_t)$. Then, the Cholesky decomposition of Σ is $\mathbb{P}$ such that $\Sigma = \mathbb{P}\mathbb{P}^\top$. To perform Monte-Carlo experiments, we take advantage of the following property:

$$\mathcal{N}_p(\boldsymbol{\mu}, \Sigma) = \mu + \mathbb{P}\mathcal{N}_p(0, I_p) \tag{2.24}$$

We can use this framework to simulate the future path of s_t. Given our sample period, a crucial point we are interested in is the following: **what would have been the position of the DEM/ITL exchange rate relative to its official lower limit on October 1st, 1992 if no realignment had occurred?**

The date we choose is rather arbitrary, but it is used here as an illustrative example. We consider different starting dates to simulate the exchange rate process until October 1st, 1992. These starting points are: August 1st, 1992, September 1st, 1992 and September 14th, 1992.

The algorithm has been initialized by setting $\mathbf{a}_0 = \mathbf{a}_t$, that is the initial position corresponds to the state vector obtained by the preceding Kalman filter maximum likelihood estimation. If we consider alternative samples, we get a vector of simulated values for the DEM/ITL exchange rate $\left(\tilde{s}_\tau^{(1)}, \ldots, \tilde{s}_\tau^{(n)}\right)$ at date τ, where τ was taken as October 1st, 1992 in our study. Consider $f(E_t[s_\tau])$ the density function of the expected exchange rate s_τ conditioned to the information available at date t. This function has been estimated by a nonparametric kernel approach using the simulated vector $\left(\tilde{s}_\tau^{(1)}, \ldots, \tilde{s}_\tau^{(n)}\right)$.

Figure **2.2** shows the density function evaluated for each starting date. It appears that the expected future domain of variation of the DEM/ITL rate varies dramatically depending on the starting date we choose. At the beginning of August 1992, only a small part of the domain is expected to go beyond the lower official limit of the DEM/ITL band on October 1st, 1992. It is remarkable to see how the situation has changed one month later. Now, there is a probability greater than $\frac{1}{2}$ for the expected DEM/ITL rate to lie outside its fluctuation margins on October 1st, 1992.

Fig. 2.2 about here

The cumulated density function can be obtained by a numerical integration[14]. Figure **2.3** depicts the resulting cumulative probability distributions which show how the

[14] We have chosen here a trapezoidal algorithm.

probability to go beyond the lower fluctuation margin increased dramatically.

Fig. 2.3 about here

The probability with which the expected value of the DEM/ITL rate on October 1st, 1992 is **below** the lower limit of its fluctuation band (s_l) can be determined. An estimate of this probability is given by:

$$\hat{p}_t = \int_{-\infty}^{s_l} \hat{f}\left(E_t\left[s_\tau\right]\right) ds_\tau \tag{2.25}$$

where $\hat{f}\left(E_t\left[s_\tau\right]\right)$ is the estimated density function based on the Gaussian kernel[15]. Given these results, we have decided to perform a recursive estimation of that probability considering starting dates ranging from May 1st, 1992 to September 11th, 1992. Figure **2.4** shows the evolution the probability for the Italian lira to lie outside its band *on October 1st, 1992. Five episodes* seem to have characterized the considered sub-period. From May 1st to June 5th, the probability $\hat{p}$ is relatively low: it fluctuates between 0.1 and 0.2. The Danish refusal to the Maastricht Treaty seem to be associated with a sudden rise in the probability (which is now above 0.2) for the DEM/ITL rate to lie outside its band on October 1st. A sharp rise in that probability intervenes in the middle of July which is exceeding 0.4. But it is rapidly decreasing until the first days of August. Then the probability grows dramatically to reach at least 0.8 just before the day of the parity shift. It is interesting to note how well this probability reflects the tensions in the DEM/ITL rate at that time. This econometric approach seems thus to give a useful indicator of the credibility of a bilateral exchange rate band.

Fig. 2.4 about here

As we have shed light on the credibility problems faced by the DEM/ITL parity few months before its revision on September 14th, 1992, we are going to investigate, in the following section, to what extent the ITL crisis can be explained by the dynamics of its DEM exchange rate.

[15] See, among others, HÄRDLE and LINTON (1994) on nonparametric techniques. Of course, the true probability should be:

$$\hat{p}_t = 1 - \int_{s_l}^{s_u} \hat{f}\left(E_t\left[s_\tau\right]\right) ds_\tau$$

but the probability to be above the upper limit s_u is here negligeable.

3. Searching for specific factors to the Italian breakdown

Two factors will be successively envisaged. First, we would like to know how the error mechanism (put in evidence in the preceding section) evolved since the narrowing of the DEM/ITL band. Next, we will analyze to what extent the DEM/ITL movements resulted from the potential growing of a speculative bubble.

3.1. Was there a cumulative depreciation?

Given the results in the section above, suppose now that:

$$s_t = s_{t-1} + \rho_t \left(s_{t-1} - s_{t-2}\right) + \varepsilon_t \tag{3.1}$$

with:

$$\rho_t = \begin{cases} \rho_t^{(1)} & \text{if } t < t^\star \\ \rho_t^{(2)} & \text{if } t \geq t^\star \end{cases} \tag{3.2}$$

such that:

$$\rho_t^{(i)} = \rho_t^{(i)} + \eta_t^{(i)} \quad i = 1, 2 \tag{3.3}$$

We suppose $\eta_t^{(i)} \sim \mathcal{N}\left(0, \sigma_i^2\right)$ for $i = 1, 2$. This implies that we allow the degree of correction ρ_t to vary over time. Moreover, we assume that the generating process of this parameter may change at a given date $t^\star$. Put in a state-space form, we obtain:

$$\begin{cases} s_t = Z_t \begin{bmatrix} \rho_t^{(1)} \\ \rho_t^{(2)} \end{bmatrix} + s_{t-1} + \varepsilon_t \\ \begin{bmatrix} \rho_t^{(1)} \\ \rho_t^{(2)} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} \rho_t^{(1)} \\ \rho_t^{(2)} \end{bmatrix} + \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} \eta_t^{(1)} \\ \eta_t^{(2)} \end{bmatrix} \end{cases} \tag{3.4}$$

with $H_t = \sigma_\varepsilon^2$ the variance of ε_t and $Q_t = \begin{bmatrix} \sigma_1^2 & 0 \\ 0 & \sigma_2^2 \end{bmatrix}$. Note that the matrix Z_t is time-dependent:

$$Z_t = \begin{cases} \begin{bmatrix} \left(s_{t-1} - s_{t-2}\right) & 0 \end{bmatrix} & \text{if } t < t^\star \\ \begin{bmatrix} 0 & \left(s_{t-1} - s_{t-2}\right) \end{bmatrix} & \text{if } t \geq t^\star \end{cases} \tag{3.5}$$

Initial values of $\rho^{(1)}$ and $\rho^{(2)}$ were set equal to zero and $P_0 = \begin{bmatrix} 0 & 0 \\ 0 & 0 \end{bmatrix}$.

Table 3

Parameter	Value	Std. error	t-Student	p-value	$\mathbf{t}_k^H$
σ_ε	0.001195	0.000033	36.21	0.00	0.169
σ_1	0.008713	0.009033	0.96	0.34	0.046
σ_2	0.003872	0.005834	0.66	0.51	0.021

Only is σ_ε statistically different from zero at the 1 percent level of risk (see column 5 of table **3**). We also have $\mathbf{t}^H = 1.0357$ and $\mathbf{t}^N = 1.0359$. Given the critical values reported in the above section, we may infer that *there is global stability of the state-space model over the whole period.* Column 6 of table **3** clearly shows that **each** parameter is also time-invariant. We represent the variations of the $\rho_t^{(i)}$ coefficients in figure **3.1**.

Fig. 3.1 about here

The dashed vertical line indicates the breaking date t^* which we fix here at June 2nd, 1992. This date corresponds to the Danish 'no' to the ratification of the Maastricht Treaty. The coefficient ρ_t seems to evolve contrary to one should have expected. From January to mid August 1990, there seems to be a correcting mechanism at work on the DEM/ITL rate. There is a brutal increase of that parameter which remains positive until the first week of June 1992 signalling a cumulating phase of DEM/ITL variations. After the Danish referendum, ρ_t diminishes quickly and becomes negative. It looks as if there were a force trying to fight against the tendency toward a depreciation of the lira *vis a vis* the mark just before the ERM crisis. This backward force started to weaken noticeably at the end of August 1992. *One possible interpretation of the time-variant nature of* ρ_t *might be that this was the resulting effect of two conflicting forces reflecting the coexistence of two groups of agents at that time*: the first one (composed perhaps with noise traders) might have pushed the lira to a further depreciation against the mark which undermined the credibility of that band, *while* the second one (reassembling agents like monetary authorities and/or smart-money investors) could have acted to preserve these fluctuation margins.

All in all, the apparition of self-cumulating movements does not seem accurate to explain the ITL crisis in September 1992.

3.2. Did a Blanchard bubble explode?

At the theoretical level, few studies investigate the consequences of the presence of bubble phenomena under an exchange rate target zone regime[16]. Recently, Ikeda and

[16] We refer the reader to the studies of Buiter and Pesenti (1990) and Miller and Weller (1990).

SHIBATA (1995) discuss the implications of bubbles on exchange rate regime switching. They conclude that variance bound tests may not be appropriate to identify bubbly dynamics. This leads us to follow the lines of WU (1995) who offers an interesting application of the Kalman filter. Unlike Wu, we do not rely here on a particular model of exchange rate determination. Following Blanchard, we suppose that exchange rate variations are submitted to a rational speculative bubble which implies the generating process:

$$\Delta s_t = b_t + \varepsilon_t \tag{3.6}$$

We assume that the bubble B_t begins to expand at a given date $t^\star$ such that:

$$b_t = \begin{cases} 0 & \text{if } t < t^\star \\ \Delta B_t & \text{if } t \geq t^\star \end{cases} \tag{3.7}$$

B_t is a bubble which grows at rate β:

$$B_t = \beta B_{t-1} + \eta_t \tag{3.8}$$

where $\eta_t \sim \mathcal{N}\left(0, \sigma_\eta^2\right)$. The corresponding state-space form can be written as:

$$\begin{cases} \Delta s_t = Z_t \begin{bmatrix} B_t \\ B_{t-1} \end{bmatrix} + \varepsilon_t \\ \begin{bmatrix} B_t \\ B_{t-1} \end{bmatrix} = \begin{bmatrix} \beta & 0 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} B_{t-1} \\ B_{t-2} \end{bmatrix} + \begin{bmatrix} 1 \\ 0 \end{bmatrix} \eta_t \end{cases} \tag{3.9}$$

with

$$Z_t = \begin{cases} \begin{bmatrix} 0 & 0 \end{bmatrix} & \text{if } t < t^\star \\ \begin{bmatrix} 1 & -1 \end{bmatrix} & \text{if } t \geq t^\star \end{cases} \tag{3.10}$$

Similarly to the preceding analysis, we have first decided to set $t^\star$ at June 2nd, 1992. However, estimates reveal that $\hat{\sigma}_\eta = 0$. No statistical inference on the presence of bubbly dynamics can be made. Given the preceding conclusions and the observation of figure **2.1**, $t^\star$ is then chosen to August 1st, 1992.

Table 4

Parameter	Value	Std. error	t-Student	p-value	$\mathbf{t}_k^H$
β	1.000436	0.104723	9.55	0.00	0.0046
σ_ε	0.001199	0.000033	36.57	0.00	0.214
σ_η	0.000012	0.000436	0.026	0.98	0.0046

Table **4** shows that β and σ_ε are statistically significant at the 1 percent level, while σ_η is not. We have $\mathbf{t}^H = 0.4784$ and $\mathbf{t}^N = 0.4784$ indicating global stability of the model. The last column of table **4** also shows individual time-invariance of parameters. Figure **3.2** illustrates the evolution of the bubble component of the DEM/ITL rate from August 1st to September 11th, 1992. The value of B_t is practically null during this period. Constructing a confidence interval at the 5 percent level of risk leads us to reject the presence of a speculative bubble *à la* Blanchard in the DEM/ITL rate.

Fig. 3.2 about here

At this point, evidence of specific factors to the Italian crisis is rather mixed. There seems to be neither a self-cumulating force at work, nor a speculative bubble. As it has been previously suggested, we did not care about a crucial feature of the ERM: its multilateral setting. This is what we are going to examine in section 4.

4. Managing the ERM as a multilateral system

Few empirical studies took explicitly the mulitlateral nature of the ERM into account[17]. We wish to answer to three main questions:

- Can we isolate a group of bilateral exchange rates (against the mark) within the ERM whose fluctuations are governed by a common source?
- Does such a group possess a hierarchic structure?
- What are the consequences for exchange rate targeting?

4.1. Identifying a common component to exchange rate fluctuations

Rose and Svensson (1994) have recently concluded that a common factor might have influenced the credibility of all DEM exchange rate bands. To follow their idea of shared credibility, we assume that log-deviations of the exchange rate s_t from its parity $s^\star$ can be decomposed into two components:

$$s_t - s^\star = \mu_t + \lambda_t \tag{4.1}$$

[17] Mundaca (1989, 1994) and Honohan (1993) offer some tentatives in this domain.

μ_t represents movements due to factors relative to the considered bilateral exchange rate while λ_t groups variations common to a predetermined set of such bilateral rates together[18]. We suppose that each component evolves as a random walk, that is:

$$\mu_t = \mu_{t-1} + \eta_t^{(\mu)} \tag{4.2}$$

and

$$\lambda_t = \lambda_{t-1} + \eta_t^{(\lambda)} \tag{4.3}$$

$\eta_t^{(\mu)}$ and $\eta_t^{(\lambda)}$ are white noise processes with respective variances σ_μ^2 and σ_λ^2.
We consider that four currencies adhere to a system of bilateral fluctuation bands. All bilateral rates are evaluated against one of these currencies which is taken as the anchor of the system. In the ERM, this role is generally devoted to the German mark. Such an exchange rate organization can be put in the following state-space form:

$$\left\{ \begin{array}{l} \begin{bmatrix} s_{1,t} - s_1^\star \\ s_{2,t} - s_2^\star \\ s_{3,t} - s_3^\star \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & \rho_1 \\ 0 & 1 & 0 & \rho_2 \\ 0 & 0 & 1 & \rho_3 \end{bmatrix} \begin{bmatrix} \mu_{1,t} \\ \mu_{2,t} \\ \mu_{3,t} \\ c_t \end{bmatrix} \\ \begin{bmatrix} \mu_{1,t} \\ \mu_{2,t} \\ \mu_{3,t} \\ c_t \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \mu_{1,t-1} \\ \mu_{2,t-1} \\ \mu_{3,t-1} \\ c_{t-1} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \eta_t^{(\mu_1)} \\ \eta_t^{(\mu_2)} \\ \eta_t^{(\mu_3)} \\ \eta_t^{(c)} \end{bmatrix} \end{array} \right. \tag{4.4}$$

Variable $\mu_{i,t}$ describes the specific part of the movements in the log-deviations of the exchange rate $s_{i,t}$ from its central parity $s_i^\star$ for $i = 1, 2, 3$. c_t is the factor common to the former exchange rates with respective contribution ρ_i. Note that these last coefficients are supposed to be constant over the entire period we examine. All stochastic terms $\eta_t^{(\mu_1)}$, $\eta_t^{(\mu_2)}$, $\eta_t^{(\mu_3)}$ and $\eta_t^{(c)}$ are white noise processes such that the covariance matrix is:

$$Q_t = \begin{bmatrix} \sigma_{\eta_1}^2 & 0 & 0 & 0 \\ 0 & \sigma_{\eta_2}^2 & 0 & 0 \\ 0 & 0 & \sigma_{\eta_3}^2 & 0 \\ 0 & 0 & 0 & \sigma_{\eta_c}^2 \end{bmatrix} \tag{4.5}$$

[18]A similar representation is also adopted by Stock and Watson (1991) to identify a common source of movements in many macroeconomic variables which is viewed by these authors as the "reference cycle".

From past empirical studies on the 1992 ERM experience[19], it appears that the Italian lira, the Spanish peseta and the British pound behaved quite differently from other ERM participating currencies. The former monies were the first to suffer from speculative attacks in September 1992 which led to the devaluation of the lira on September 14th. Few days after this parity adjustment, the lira and the sterling abandoned the ERM, while the peseta was devaluated against all remaining ERM members. That the speculative attack concentrated on these three currencies invited us to investigate the potential existence of a common source of their respective price variation against the mark.

For this purpose, we have run the above estimating procedure to identify this common factor. Figures **4.1** to **4.3** draw the results obtained by Kalman filtering for the Italian, Spanish, and British cases respectively. In the upper part of each graphic, we draw the observed exchange rates deviations from their respective central parity $(s_t - s^\star)$ during the sample period. In the lower part of the following figures, we show the decomposition of these movements into the individual (μ_t) and the shared (c_t) components.

- We are led to conclude that a common, but unobservable, factor seems to be effectively at work during the whole period. It is remarkable to observe that large **movements** in the three DEM exchange rates are explained by this *shared* component while the *individual* factor determines only the exchange rate **level**. This is particularly the case for the peseta (figure **4.2**), and, to a lesser extent, for the lira (figure **4.1**). At the opposite, most of DEM/GBP variations seem to have been due to their own country developments: figure **4.3** reveals this feature.

- During the month preceding the September 1992 crisis, the rapid depreciation of the three currencies relative to the mark seems to constitute a singular episode of the ERM functioning. From the figures to **4.1** and **4.2**, we observe that the trend in exchange rates is essentially determined by the common, though unobservable, part in the cases of the peseta and the lira. On the contrary, the rapid decline of the DEM/GBP rate resulted from bilateral conditions between Germany and the United Kingdom as figure **4.3** illustrates it.

- Although not reported here, similar estimations with the (ITL, ESP) pair and another participating currency show that only ITL, ESP, and GBP seem to share the same common factor.

[19] Seee, among others, Eichengreen and Wyplosz (1993), Rose and Svensson (1994).

To conclude, it appears that movements in the DEM/ITL and DEM/ESP rates were determined most of the part jointly, even in the few months preceding the 1992 crisis. At the same time, only specific factors explained most of the DEM/GBP rate variations. We underline the fact that these former factors should be viewed here as being specific in the sense that there were not shared by the other two currencies.

Fig. 4.1 about here

Fig. 4.2 about here

Fig. 4.3 about here

4.2. Further evidence on causal relationships between ERM currencies

We have just shown above that a common factor seems to have influenced the evolutions of the lira, the peseta and the sterling against the German mark. It seems interesting to examine more profoundly the nature of the interactions between theses three bilateral exchange rates. Gerlach and Smets (1994) have built a VAR model to shed light on contagious phenomena in exchange rate crises. Such a representation proves to be appropriate to search for causal relationships among exchange rate returns. In the first paragraph, we are going to expose the principle of a VAR estimation. In the second paragraph we will apply it to our trivariate exchange rate system.

4.2.1. The VAR estimation methodology

$\{y_t\}$ is a VAR process of order p with dimension K if we have:

$$\begin{cases} y_t = v + A_1 y_{t-1} + \ldots + A_p y_{t-p} + u_t \qquad t = 1, \ldots, T \\ u_t \sim \mathcal{N}(0, \Sigma_u) \end{cases} \tag{4.6}$$

and

$$y_t = \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ \vdots \\ y_{K,t} \end{bmatrix} \tag{4.7}$$

We use LÜTKEPOHL's (1991) results (see page 63 of his book). The above VAR(p) process can be written as[20]:

$$Y = BZ + u \tag{4.8}$$

Let $y = \operatorname{vec} Y$ and $\beta = \operatorname{vec} B$. Lütkepohl establishes the subsequent relation:

$$\hat{\beta} = \left[\left(ZZ^{\top} \right)^{-1} Z \otimes I_K \right] y \tag{4.9}$$

$\hat{\beta}$ is the maximum likelihood estimator of β. The covariance matrices of estimators are then defined by:

$$V\left[\hat{\beta}\right] = \left(ZZ^{\top} \right)^{-1} \otimes \hat{\Sigma}_u \tag{4.10}$$

noting that

$$\hat{\Sigma}_u = \frac{1}{T} Y \left[I_T - Z^{\top} \left(ZZ^{\top} \right)^{-1} Z \right] Y^{\top} \tag{4.11}$$

Let $\mathbb{P}$ the Cholesky decomposition of Σ_u (that is $\Sigma_u = \mathbb{P}\mathbb{P}^{\top}$) and $\mathbb{P}^{*} = \operatorname{vech} \mathbb{P}$. Then we verify:

$$\hat{\Sigma}_u = \overset{\wedge}{\mathbb{P}} \overset{\wedge}{\mathbb{P}}^{\top} \tag{4.12}$$

$$\overset{\wedge}{\mathbb{P}}^{*} = \operatorname{vech} \overset{\wedge}{\mathbb{P}} \tag{4.13}$$

with

$$V\left[\overset{\wedge}{\mathbb{P}}^{*}\right] = H \frac{2\mathbf{D}_K^{+} \left(\hat{\Sigma}_u \otimes \hat{\Sigma}_u \right) \mathbf{D}_K^{+\top}}{T} H^{\top} \tag{4.14}$$

where $\mathbf{D}_K^{+}$ is the Moore-Penrose inverse of the duplication matrix $\mathbf{D}_K$ and

$$H = \left[\mathbf{L}_K \left(I_{K^2} + \mathbf{K}_{KK} \right) \left(\overset{\wedge}{\mathbb{P}} \otimes I_K \right) \mathbf{L}_K^{\top} \right]^{-1} \tag{4.15}$$

$\mathbf{L}$ et $\mathbf{K}$ represent the elimination and commutation matrices respectively.

[20] We use the following notations:

$$\begin{aligned}
Y &= \begin{bmatrix} y_1 & y_2 & \cdots & y_T \end{bmatrix} \qquad B = \begin{bmatrix} v & A_1 & A_2 & \cdots & A_p \end{bmatrix} \\
Z_t &= \begin{bmatrix} 1 \\ y_t \\ \vdots \\ y_{t-p+1} \end{bmatrix} \\
Z &= \begin{bmatrix} Z_0 & Z_1 & \cdots & Z_{T-1} \end{bmatrix} \qquad U = \begin{bmatrix} u_1 & u_2 & \cdots & u_T \end{bmatrix} \\
u &= \operatorname{vec} U
\end{aligned}$$

4.2.2. A trivariate analysis

We have shown in paragraph 1 of section 3 above that there was evidence of a noticeable change in the driving process of the DEM/ITL rate since June 2nd,1992. Therefore, we decide to split our sample into two sub-periods. The first one extends from October 8th, 1990 (the date of the GBP entry into the ERM) to June 1st, 1992, that is just before the announcement of the Danish 'Nej' to the ratification of the Maastricht Treaty. The second period corresponds to the months just preceding the September turbulences within the ERM[21].

We have built **a trivariate VAR** model composed with exchange rate **returns**, that is $s_t - s_{t-1}$.

Let $y_t = \begin{bmatrix} s_{1,t} - s_{1,t-1} \\ s_{2,t} - s_{2,t-1} \\ s_{3,t} - s_{3,t-1} \end{bmatrix}$ such that the number in index refers to ITL (1), ESP (2), and GBP (3). The first step consists in determining the optimal lag of the VAR process. We exploit the results derived by Deniau, Fiori and Mathis (1992) on this particular point.

Table 5

	BIC	AICc	SIC	FPE	AIC	HQ
1992:06:02-1992:09:13	1	1	1	2	3	1

To set the maximal lag order of the VAR model, we rely on short criteria in table **5**. Retaining one lag seems to be adequate.

[21] Due to space constraints, we do not report here the results for the period preceding June 1992. We will only signal them as a basis of comparison.

In the next step we estimate the VAR process by the maximum likelihood method. This leads to the representation below:

$$\begin{bmatrix} s_{1,t} - s_{1,t-1} \\ s_{2,t} - s_{2,t-1} \\ s_{3,t} - s_{3,t-1} \end{bmatrix} = \begin{bmatrix} -0.2173 & 0.0702 & 0.0847 \\ 0.3984^{*} & -0.0955 & -0.0098 \\ -0.5534^{*} & 0.1826 & -0.1777 \end{bmatrix} \begin{bmatrix} s_{1,t-1} - s_{1,t-2} \\ s_{2,t-1} - s_{2,t-2} \\ s_{3,t-1} - s_{3,t-2} \end{bmatrix} + \begin{bmatrix} -0.0001 \\ -0.0005^{*} \\ -0.0008^{*} \end{bmatrix} \tag{4.16}$$

An asterisk denotes the statistical significance of the corresponding coefficient at the 5 percent risk level. A striking result of the estimation above is the lack of influence of DEM/ESP (DEM/GBP respectively) returns on DEM/ITL and DEM/GBP (DEM/ITL and DEM/ESP resp.) ones. Only does the DEM/ITL rate of variation influence the other two returns from June 2nd, 1992 to September 11th, 1992. *This suggests that the lira had a leading role in the occurrence of the September 1992 crisis.* One should have in mind that running a similar estimation over the first sub-period (October 1990-June 1992) leads to reject the statistical significance of any linkage between these three returns. This suggests a completely different functioning of the ERM during this period of greater tranquillity.

We were interested in studying *Granger causality* relationships among the three DEM exchange rates. For this purpose, tests based on Wald's principle have been performed. These new tests are based on an estimate of the covariance matrix, unlike previous Wald tests performed in section 2 which rested on the information matrix. Thus, expression (2.19) is now replaced by:

$$\mathbf{W} = R\left(\hat{\theta}\right)^{\top} \left[\frac{\partial R}{\partial \theta^{T}}\left(\hat{\theta}\right) V\left[\hat{\theta}\right] \frac{\partial R^{T}}{\partial \theta}\left(\hat{\theta}\right)\right]^{-1} R\left(\hat{\theta}\right) \tag{4.17}$$

Before interpreting the results, we need first to define what the non causality hypothesis means.

Granger causality:

Consider our VAR(1) process. Variable j does not cause in the Granger sense the set of variables I if

$$\beta_{3(i+1)+j} = 0 \quad \text{for all } i \in I$$

Estimations lead to the following results reported in table **6** where $\not\rightarrow$ means "does not cause in the Granger sense":

Table 6

1992:06:02-1992:09:13	**W**	p-value
ITL $\not\rightarrow$ ESP	4.2854	0.038
ITL $\not\rightarrow$ GBP	4.3152	0.038
ITL $\not\rightarrow$ (ESP,GBP)	17.0719	0.000
ESP $\not\rightarrow$ ITL	0.4091	0.522
ESP $\not\rightarrow$ GBP	0.7644	0.382
ESP $\not\rightarrow$ (ITL,GBP)	1.0177	0.601
GBP $\not\rightarrow$ ITL	1.5317	0.216
GBP $\not\rightarrow$ ESP	0.0108	0.917
GBP $\not\rightarrow$ (ITL,ESP)	2.5683	0.277

We observe the prominent influence of the DEM/ITL rate upon the other two exchange rates. The null hypothesis of non causality from the DEM/ITL rate to DEM/ESP and/or DEM/GBP returns is unambiguously rejected. At the opposite, there is no evidence of causal relations from neither the DEM/ESP nor the DEM/GBP returns to whatever set of the remaining returns considered. Besides, we wish to know if the couple (DEM/ESP,DEM/GBP) does not Granger cause the DEM/ITL return over this sub-period. In this case, the null hypothesis is:

$$\begin{cases} \beta_7 = 0 \\ \beta_{10} = 0 \end{cases} \tag{4.18}$$

The corresponding value of the Wald statistic is 2.8932 and its p-value equals 0.235. *This clearly indicates the absence of any causal link (in the Granger sense) from (DEM/ESP,DEM/GBP) to the DEM/ITL return.* These relations can be opposed to the ones found for the first sub-period during which there is no evidence of causal relationships from one return to another.

Given these conclusions, we were also interested in identifying **instantaneaous causality** between exchange rate returns. We define instantaneous causality as follows:

Consider a set I of variables extracted from a VAR model. Testing instantaneous non causality among variables belonging to I reduces to check the validity of the subsequent equality:

$$\mathbb{P}^*_{\frac{i'(i'-1)}{2}+i''} = 0 \quad \text{for all } (i', i'') \in I \text{ such that } i' > i'' \tag{4.19}$$

Table 7

1992:06:02-1992:09:13	**W**	p-value
(ITL,ESP)	29.6433	0.000
(ITL,GBP)	2.1764	0.140
(ESP,GBP)	16.5889	0.000

From the table above, *we cannot reject the null of no instantaneous causality only for the* (ITL,GBP) *case.* Applying the same test to the sub-period of reference, we find significant instantaneous causality between each pair of exchange rate returns considered above, contrasting with the absence of Granger-causality.

Having shown evidence of causal relations between exchange rate returns, it seems interesting to build a structural VAR model. This enables to take some restrictions into consideration. In particular, we take into account the fact that:

- the DEM/ESP return does not Granger-cause the DEM/GBP and DEM/ITL returns,
- the DEM/GBP return does not Granger-cause both DEM/ITL and DEM/ESP returns,
- there is no instantaneous causal relationship between the DEM/ITL and DEM/GBP returns.

All these constraints lead to the following state-space form of the structural VAR model:

$$\left\{ \begin{array}{l} \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} \\ \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} = \begin{bmatrix} \theta_4 & 0 & 0 \\ \theta_5 & \theta_7 & 0 \\ \theta_6 & 0 & \theta_8 \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ y_{2,t-1} \\ y_{3,t-1} \end{bmatrix} + \begin{bmatrix} \theta_1 \\ \theta_2 \\ \theta_3 \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} u_{1,t} \\ u_{2,t} \\ u_{3,t} \end{bmatrix} \end{array} \right. \tag{4.20}$$

where $u_{1,t}, u_{2,t}, u_{3,t}$ is a white noise process with covariance matrix Q such that:

$$Q = \Sigma_u = \mathbb{P}\mathbb{P}^\top \tag{4.21}$$

$\mathbb{P}$ stands for the Cholesky decomposition matrix of Σ_u. Let $\mathbb{P}^*$ defined by:

$$\mathbb{P}^* = \operatorname{vech} \mathbb{P} \tag{4.22}$$

Note that the above formulation is time-invariant: matrices Z_t, $\mathbf{d}_t$, H_t, $\mathbf{T}_t$, $\mathbf{c}_t$, $\mathbf{R}_t$ and Q_t do not depend on time. That is why the time subscript t is omitted here. The identified constraints on instantaneous causality between exchange rate returns implies that $\mathbb{P}^*$ can be expressed as:

$$\mathbb{P}^* = \begin{bmatrix} \theta_9 \\ \theta_{10} \\ 0 \\ \theta_{11} \\ \theta_{12} \\ \theta_{13} \end{bmatrix} \tag{4.23}$$

The application of the Kalman filter maximum likelihood procedure on the second sub-period gives[22]:

$$\begin{cases} \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} \\ \begin{bmatrix} y_{1,t} \\ y_{2,t} \\ y_{3,t} \end{bmatrix} = \begin{bmatrix} -0.1392 & 0 & 0 \\ 0.4856^{**} & -0.2245^{*} & 0 \\ -0.4560^{*} & 0 & -0.0923 \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ y_{2,t-1} \\ y_{3,t-1} \end{bmatrix} + \begin{bmatrix} -0.0002 \\ -0.0005^{*} \\ -0.0009^{**} \end{bmatrix} \end{cases} \tag{4.24}$$

and

$$\hat{\mathbb{P}}^* = \begin{bmatrix} 1.353^{**} \\ 0.942^{**} \\ 0 \\ 1.494^{**} \\ 1.300^{**} \\ 2.177^{**} \end{bmatrix} \times 10^{-3} \tag{4.25}$$

Almost all estimated parameters are significantly different from zero at the 5 percent (*) or at the 1 percent (**) level of risk. From the comparison of the unrestricted VAR model with the above constrained version, we conclude that parameter signs are preserved, while noticeable changes in magnitude can observed. A rise in the DEM/ITL return still induces a fall in the DEM/GBP return and a rise in DEM/ESP one. But,

[22] No similar estimation has been performed for the first sub-interval given the lack of constraints.

in absolute value, the relative impact is now reversed. It is yet $\frac{0.4856}{0.4560} = 1.06$, instead of being $\frac{0.3984}{0.5534} = 0.72$ in the unrestricted formulation of the VAR.

Due to the time-invariance nature of the preceding state-space representation, it is interesting to examine some useful properties of the latter. In particular, *the model is stable*. This means that the modulus of the eigenvalues associated with the transition matrix T are all less than one[23]. The crucial implication of this finding is that the error-covariance matrix is time-invariant in the steady-state or equivalently:

$$\lim_{t \to +\infty} P_{t|t-1} = \bar{P} \tag{4.26}$$

The matrix $\overline{P}$ is in fact the solution of the algebraic Riccati equation below[24]:

$$\bar{P} - T\bar{P}T^{\top} + T\bar{P}Z^{\top} \left[Z\bar{P}Z^{\top} + H \right]^{-1} Z\bar{P}T^{\top} - RQR^{\top} = \mathbf{0}_{3\times3} \tag{4.27}$$

where $\mathbf{0}_{3\times3}$ is a 3×3 matrix of zeros. Equation (4.27) can be solved numerically[25]. Our computations give:

$$\bar{P} = \begin{bmatrix} 1.8307 & 1.2748 & 0.000 \\ 1.2748 & 3.1197 & 1.9415 \\ 0.0000 & 1.9416 & 6.2497 \end{bmatrix} \times 10^{-6} \tag{4.28}$$

The steady-state error correlation matrix Ω is then:

$$\Omega = \begin{bmatrix} 1 & 0.53 & 0.00 \\ 0.53 & 1 & 0.43 \\ 0.00 & 0.43 & 1 \end{bmatrix} \tag{4.29}$$

In the steady-state, there would exist a positive correlation between the DEM/ITL and DEM/ESP returns on one hand and, on the other hand between the DEM/ESP and DEM/GBP pair. Thus, a further depreciation of the Italian (Spanish respectively) currency against the mark is associated with a further depreciation of the peseta (British pound respectively) *vis à vis* the German mark. We also wish to underline the absence of any link between the returns of the lira and the sterling against the mark in the steady-state. This last observation confirms our preceding findings on causal relationships between these exchange rate returns.

[23] The three characteristic roots are real with respective values: -0.0923, -0.224, -0.139.

[24] See Harvey (1989) on properties of time-invariant state-space models, especially pages 113 to 125.

[25] To this end, we have used Aoki's (1987) algorithm (see page 100 of his book).

A related implication of the model stability is that:

$$\lim_{t \to +\infty} \mathbf{a}_t = \bar{\mathbf{a}} \tag{4.30}$$

In our case, the following expression can derived:

$$\bar{\mathbf{a}} = [I - T]^{-1} \cdot \mathbf{c} \tag{4.31}$$

with $\mathbf{c}$ the estimate of $\mathbf{c}_t$. Given the estimates, this leads to:

$$\bar{\mathbf{a}} = \begin{bmatrix} -1.462 \\ -5.052 \\ -7.313 \end{bmatrix} \times 10^{-4} \tag{4.32}$$

The long-run solution shows the inherent depreciating tendency of the peseta, the lira and the pound against the German currency: all elements of $\bar{\mathbf{a}}$ are negative. However, this last conclusion should be tempered by the fact that standard-errors lead to reject the significance of each coefficient.

4.3. Balancing costs of intervention with precision of targeting

Given the time-invariance property of the above state-space model, it appears useful to verify if the system is **controllable**. Consider a system characterized by a transition equation of the form:

$$\boldsymbol{\alpha}_t = T\boldsymbol{\alpha}_{t-1} + \mathbf{c} + \Gamma u_t \tag{4.33}$$

The controllability of this system requires that the controllability matrix $\begin{bmatrix} \Gamma & T\Gamma & \cdots & T^{m-1}\Gamma \end{bmatrix}$ is of rank m, *i.e.* the dimension of the state vector.
The fact that a system is controllable means that we can choose a set of control vectors $u = \{u_{\tau+1}, u_{\tau+2}, \ldots, u_{\tau+l}\}$ such that the state vector reaches a predetermined value $\boldsymbol{\alpha}^*$ at date $(\tau + l)$ given that the position of the system is $\boldsymbol{\alpha}_\tau$ at time τ. Such a property has meaningful implications for exchange rate policy-making. In our case, the vector u can be thought as a set of instruments which monetary authorities can choose to intervene in the foreign exchange market. In other words, controllability means here that it is possible to find an appropriate set of actions, defining an intervention strategy, in order to move *simultaneously* the three DEM/ITL, DEM/ESP and DEM/GBP rates back to their respective central parity at a known horizon. This is precisely one of the two options of the monetary authorities intervention rule suggested by Bertola and Caballero (1992) when the exchange rate hits one the limits of its band.
In our case, Γ corresponds to $R\mathbb{P}$ where $\mathbb{P}$ is the Cholesky decomposition matrix of Σ_u.

Since the singular values of the controllability matrix are all different from zero[26], *our model is controllable.*

Suppose that, at date τ, monetary authorities wish to move bilateral exchange rates against the mark at their corresponding central parity at date $(\tau + l)$, that is:

$$s_{i,\tau+l} = s_i^* \quad \text{for all } i = 1, 2, 3 \tag{4.34}$$

i symbolizes the DEM/ITL (1), DEM/ESP (2) and DEM/GBP (3) exchange rates. In the model, $s_{i,\tau+l}$ is equal to:

$$s_{i,\tau+l} = s_{i,\tau} + \sum_{j=1}^{l} \alpha_{j,\tau+j} \tag{4.35}$$

where $\alpha_{j,\tau+j}$ is the j-th element of the vector $\alpha_{\tau+j}$. Let f the application associated with the system (4.34) defined by:

$$\begin{aligned} f:\ \mathbb{R}^{3l} &\longrightarrow \mathbb{R}^3 \\ u &\longmapsto f(u) = s_{\tau+l} \end{aligned} \tag{4.36}$$

s and s^* represent the vectors of current exchange rates and parities, respectively. If we assume that there exists a solution to $f(u) = s^*$, then $\dim \operatorname{Im} f = 3$. So $\dim \ker f = 3(l-1)$. In the case when $3(l-1) > 0$, several solutions may exist. They then corresponds to the linear variety of direction $\ker f$.

There is an obvious solution to our problem. It is $u = \{\tilde{u}, \mathbf{0}_3, \ldots, \mathbf{0}_3\}$ such that:

$$\begin{aligned} \tilde{u} = \ & \left[I + T + \cdots + T^{l-1}\right]^{-1} \Gamma^{-1} \left\{(s^* - s_\tau) - \left[T + \cdots + T^l\right](s_\tau - s_{\tau-1})\right\} \\ & - \left[I + T + \cdots + T^{l-1}\right]^{-1} \Gamma^{-1} \left\{\left[I + T + \cdots + T^{l-1}\right] \mathbf{c}\right\} \end{aligned} \tag{4.37}$$

We have fixed τ as September 11th, 1992. Because the ERM prescribes that exchange rates cannot move outside pre-announced limits, solutions to the problem must be also consistent with the European institutional monetary framework.

Consider first the simplest case $l = 1$. We have a unique solution which is:

$$\tilde{u} = \begin{pmatrix} 19.243 \\ -10.159 \\ 33.391 \end{pmatrix} \tag{4.38}$$

To be able to understand what $\tilde{u}$ means, imagine that there are shocks, distributed as a three dimensional standard normal law, which can take the values in (4.38) so

[26] There are equal to 0.0027, 0.0018 and 0.0011.

as to bring the three exchange rates back to their respective parities within a day. It is impressive how high is the amount of action necessary to influence the three DEM exchange rates in a desired way. Another striking feature of this last result is that, **to defend a group of depreciated currencies against the mark, it may be appropriate to depreciate furthermore one of them** (see the negative value obtained for the second element of $\tilde{u}$). Although unappealing at first sight, this conclusion can be explained by the fact that at that time (*i.e.* on September 11th, 1992) there was still considerable room for a depreciation of the peseta *vis à vis* the mark, while the other two currencies were dangerously close to their limit.

For all $l \geq 2$, several solutions exist. To illustrate this point, we report two possible set of actions so that the three exchange rates come back to their respective parities within 3 days (or $l = 3$):

$$u = \begin{bmatrix} 5.112 & 1.943 & 13.286 \\ \text{-3.763} \;, & \text{-3.053} \;, & \text{-6.161} \\ 13.257 & 12.527 & 13.527 \end{bmatrix} \text{ and } u = \begin{bmatrix} 4.982 & 2.121 & 13.246 \\ \text{-4.310} \;, & \text{-3.500} \;, & \text{-5.399} \\ 12.754 & 13.118 & 13.560 \end{bmatrix} \tag{4.39}$$

Even if monetary authorities base their decisions on a much longer horizon (say 9 days), the required sequence of impulsions still remains surprisingly high.

However, it is important to note that we have obtained sets of control variables which cannot be compared with each other on the basis of the cost they imply. Central banks have to support costs when intervening in the foreign market. Suppose that such costs increase with the amount of intervention, or the intensity of the action u. Assume also that monetary authorities wish to minimize the cost of their interventions. If we consider a quadratic cost function, this objective can be attained by resolving the following optimization program:

$$\begin{cases} \min\limits_{u} \|u\|_2 \\ \text{s.t.} \quad s_{\tau+l} = s^* \end{cases} \tag{4.40}$$

The problem of multiple paths toward the final objective, that is the return to the middle of the band, still remains. The resulted strategies become obviously less and less costly than the previous ones as monetary authorities attach an increasing importance to the costs they support. However, this is done at the expense of reaching a value close to the parity.

We suppose here a quadratic objective function g of the form:

$$g = \|s_{\tau+l} - s^*\|_2 + \frac{1}{\lambda}\|u\|_2 \tag{4.41}$$

This objective function is similar to that suggested by MILLER and ZHANG (1995) in order to determine optimal exchange rate fluctuation margins. Like these authors, we assume that monetary policy-makers face proportional costs of intervention.

u_1 and u_2 corresponds to the results obtained for $\lambda_1 = 10^6$ and $\lambda_2 = 10^4$ (respectively), taking $l = 3$:

$$u_1 = \begin{bmatrix} 3.113 & 1.130 & 8.400 \\ \text{-0.780} \;, & \text{-0.431} \;, & \text{-1.640} \\ 9.763 & 9.672 & 10.656 \end{bmatrix} \text{ and } u_2 = \begin{bmatrix} 0.072 & \text{-0.007} & 0.358 \\ 0.632 \;, & 0.625 \;, & 0.691 \\ 1.038 & 1.028 & 1.133 \end{bmatrix} \quad (4.42)$$

The multiplicity of intervention strategies results from the specification of the cost function. We have postulated costs *proportional* to the amount of intervention. AVESANI (1990) and AVESANI and JEANBLANC-PICQUÉ (1992) show that, in an infinite horizon continuous-time framework, the optimal policy reduces to a reflection policy which consists in infinitesimal intervention at the limits of the exchange rate fluctuation band. In our case, we derive a set of **discrete interventions** so that, to FLOOD and GARBER's (1992) words, a *"particular target zone can be supported by an infinity of intervention strategies"*. By comparison with (4.39), the introduction of cost considerations implies lower impulses than before. The resulting sequences of innovations also look very different depending on the relative importance attached by monetary policy-makers to the costs. But it has to be mentioned that:

$$\left(\frac{\hat{s}_{\tau+3} - s^*}{s^*}\right)_1 = \begin{bmatrix} 0.49 \\ 1.2 \\ -0.08 \end{bmatrix} \times 10^{-2} \text{ and } \left(\frac{\hat{s}_{\tau+3} - s^*}{s^*}\right)_2 = \begin{bmatrix} 1.3 \\ -0.23 \\ -0.9 \end{bmatrix} \times 10^{-2} \quad (4.43)$$

This implies that $\|s_{\tau+l} - s^*\|_2$ is **twenty times lower** when $\lambda = \lambda_1$ than when $\lambda = \lambda_2$. *There is a trade-off between the costs of the chosen strategy and the precision with which the exchange rate will be targeted.* Balancing precision in exchange rate targeting with intervention costs may explain why monetary authorities prefer to announce fluctuation margins instead of a fixed parity. Of course, this is one possible explanation of the existence of exchange rate bands, but alternative explanations might aslo be valid[27]. Our results seem to be *quite* in accordance with Miller and Zhang's conclusions. A closer look at (4.43) reveals in fact that the relationship between precision in targeting

[27] See the literature on cheap-talk incentives of central bankers. STEIN (1989) analyses the consequences of the existing information asymmetry between market participants and policy-makers on exchange rate targeting. Such an asymmetry is also examined by AVESANI and al. (1995) in an exchange rate target zone framework.

and costs is not necessarily a monotonically decreasing one. As monetary authorities place an increasing weight on her costs, the "controlled" exchange rate lies farther away from its targeted value *only* for two of the three bilateral exchange rates. This appears to be a singular feature of multilateral exchange rate target zone regimes which traditional theoretical approaches fail to take into account.

5. Conclusions and directions for further research

To conclude, we tried, in this paper, to shed a new light on the September 1992 events.

We first propose an alternative procedure to evaluate the credibility of an exchange rate band. Since DEM/ITL fluctuations in its narrow band are well-fitted by an ARMA(2,2) process, simulations based on the Kalman filter reveal that tensions started to be particularly important since mid-August 1992. At that time, we observe a sharp increase in the probability to expect the DEM/ITL rate to go through its depreciation limit on October 1st, 1992. This result is clearly incompatible with a perfectly credible exchange rate target zone.

We then investigate potential explanations of the occurrence of the DEM/ITL realignment on September 14th, 1992. We are not able to find any self-reinforcing mechanism in the DEM/ITL process during the few months before the September 1992 exchange rate turmoil, nor there is evidence that the DEM/ITL dynamics followed a speculative bubble path *à la* Blanchard.

Given the absence of specific factors to the ITL crisis, we adopt another perspective, considering the multilateral aspect of the ERM. Distinguishing between an individual and a common component in bilateral exchange rates against the mark, we show evidence on a common, but directly unobservable, element which governed the movements of the lira, the peseta and the sterling against the mark.

Our VAR analysis underline the leading role of DEM/ITL variations during the period which preceded the September 1992 crisis. A examination of causal relationships corroborate that relations between these three exchange rates were markedly hierarchic.

Finally, we derive results which may be useful for optimal exchange rate policy-making, suggesting that there is a trade-off between the precision with which exchange rates can be targeted and the costs implied by the required actions.

To our minds, future work may be oriented into two directions. At a theoretical level, there is no detailed treatment of the functioning of a multilateral target zone system. In particular, standard models do not explain what happens when monetary authorities have to defend more than a unique bilateral band. Does she attach the same importance to all bilateral bands she cares? Given such a setting, we may be able to identify propagation mechanism of exchange rate crises. In this view, Buiter and al. (1995) offer a very promising theoretical approach. At an empirical level, it remains to be seen what role the dollar played in ERM crises. One could then follow the lines of Giavazzi and Giovannini (1985) and of Kaufmann (1985). It would be also of great interest to work with very high frequency data like hourly quotations.

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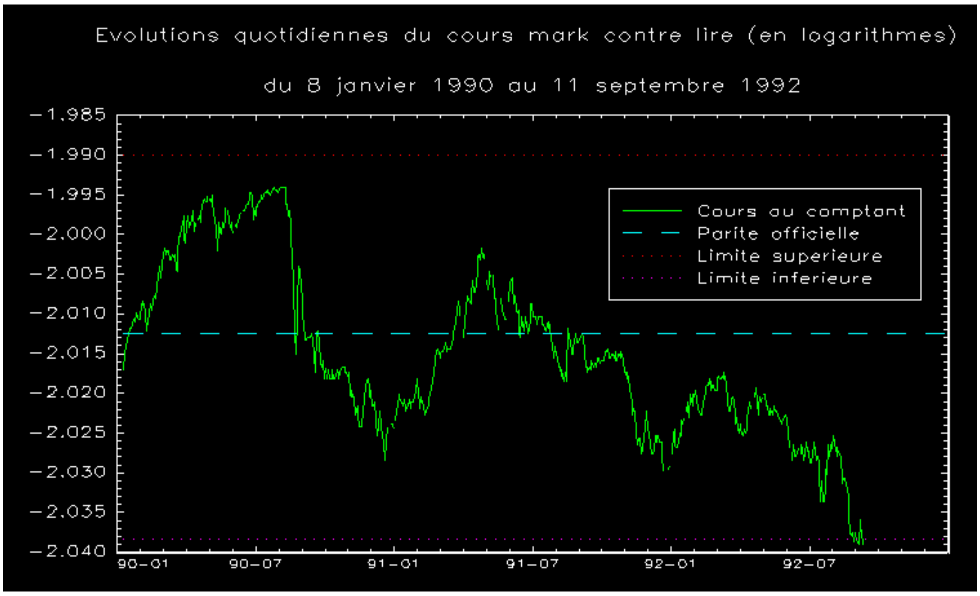

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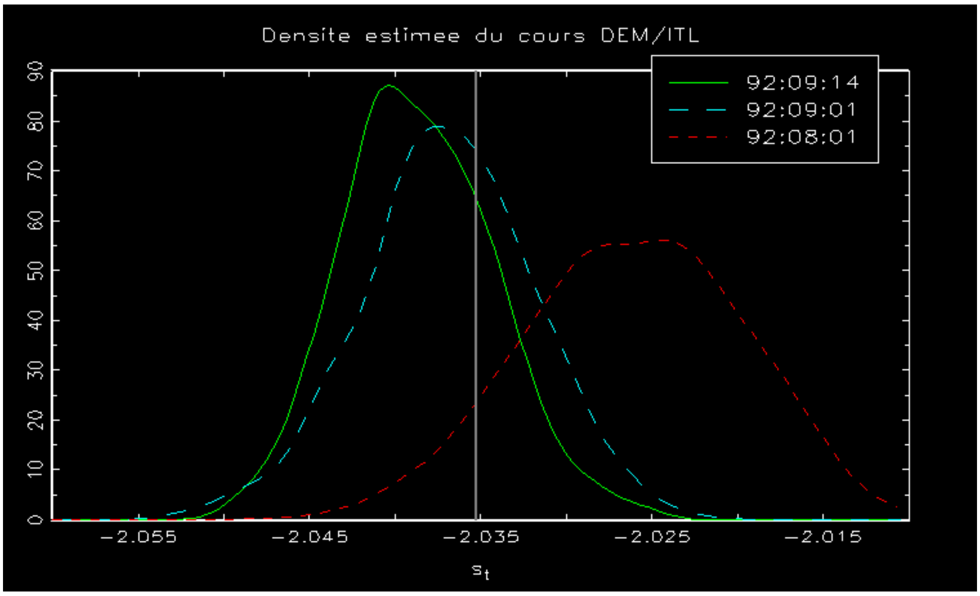
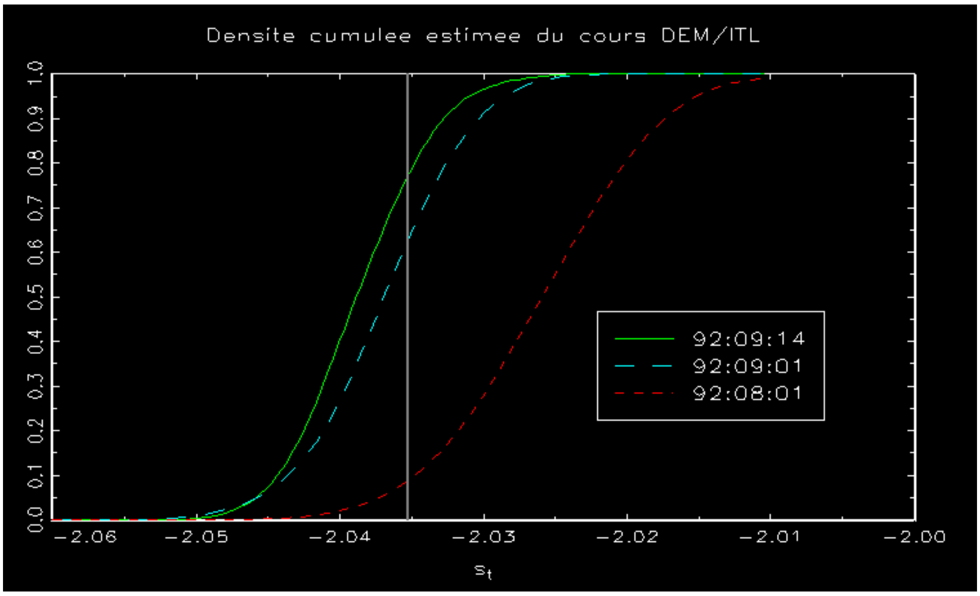


Figure 6.2:



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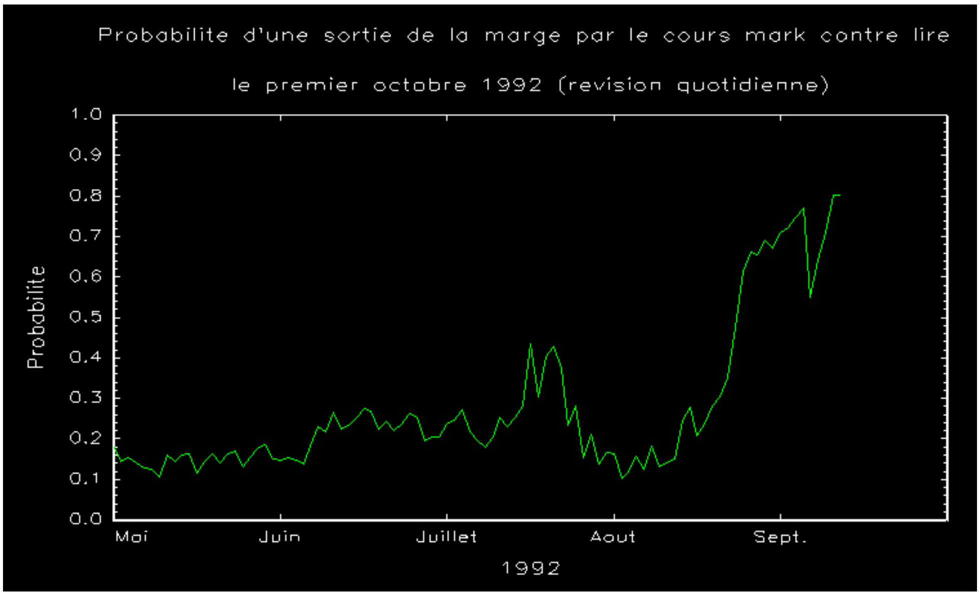

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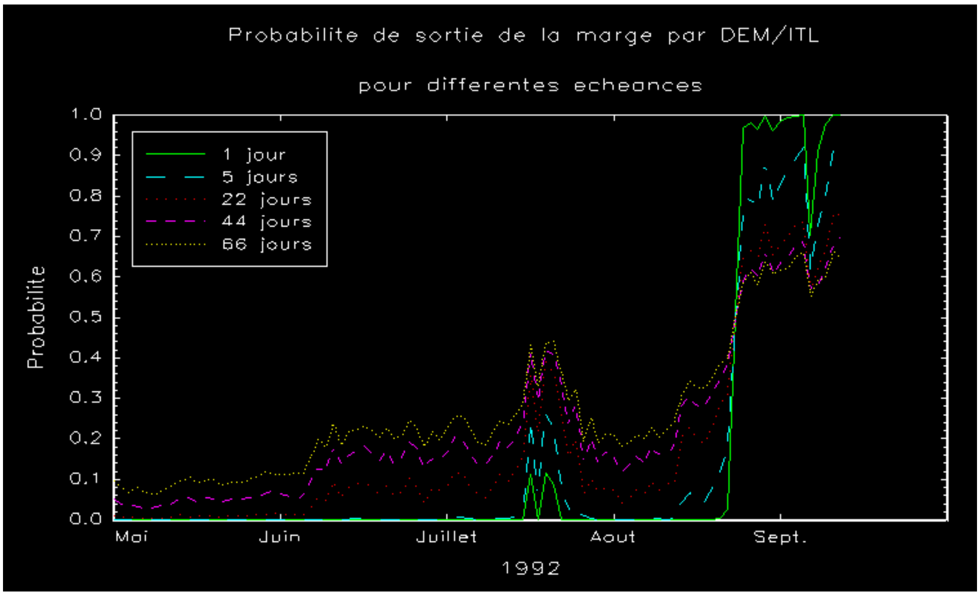

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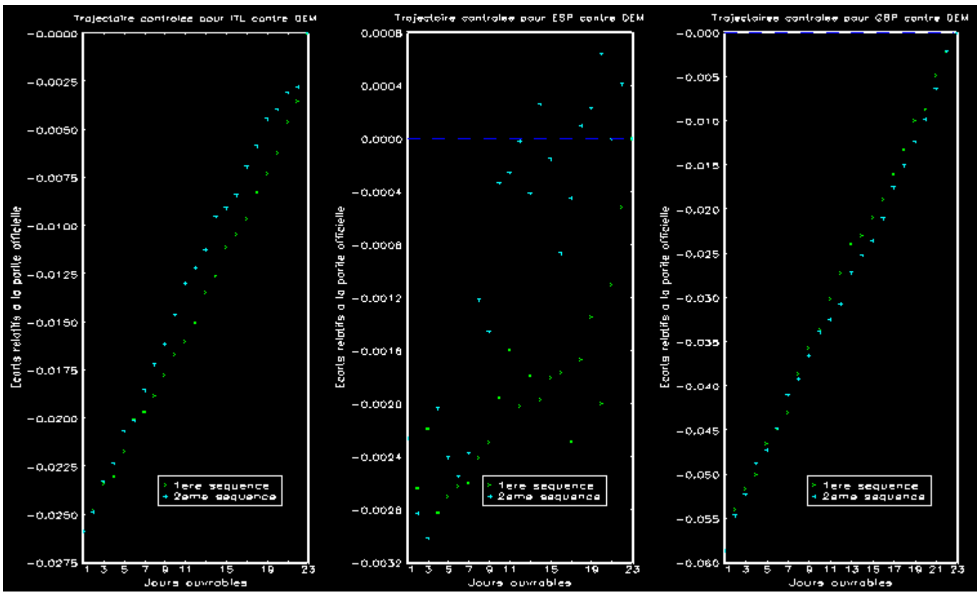

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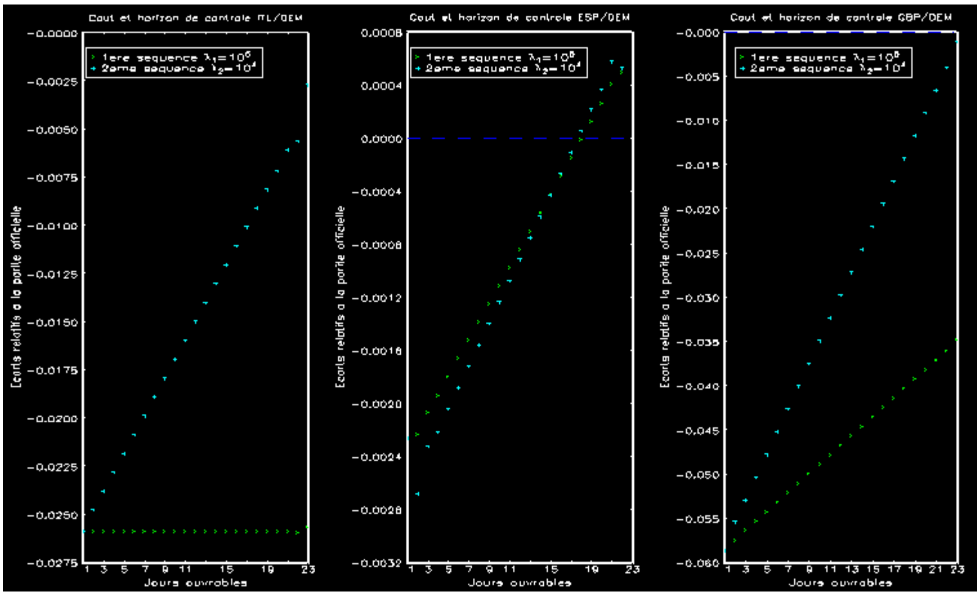
Cout et horizon de controle ITL/DEM
Cout et horizon de controle ESP/DEM
Cout et horizon de controle GBP/DEM
1ere sequence $\lambda_1=10^6$
2eme sequence $\lambda_2=10^4$
Ecarts relatifs a la parite officielle
Jours ouvrables

Figure 6.7: