Regimen changes and duration in the European Monetary System

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ABSTRACT:

This paper examines the regime changes in the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS), applying the duration model approach to weekly data of eight currencies participating in the ERM, covering the complete EMS history. When using the non-parametric (univariate) analysis, we found that for those regimens with long durations, the ERM would have been relatively stable, while for the (more common) regimes associated with short durations would have been more unstable. The probability of maintaining a certain regime is estimated to be 0.685. When applying a parametric (multivariate) analysis to investigate the role of other variables in the probability of a regime change, we conclude that the interest rate differential with Germany and the magnitude of the realignment would have negatively affected the duration of a given regime, while credibility would have positively influenced such duration. Finally, when distinguishing between groups of currencies, we observe that those in the core are more stable than those in the periphery, obtaining evidence against equality of survival functions among these groups of currencies.

JEL Codes: C41, F31, F33
Key Words: Duration models, exchange rates, European Monetary System

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1. Introduction

The accession of the Central and Eastern European countries (CEEC) as well as Malta and Cyprus presents the European Union (EU) with one of the greatest challenges in its history. In June 2001, the Gothenburg European Council clearly stated that the process of European enlargement is irreversible. The new EU member states also commit themselves to adopt the euro at a later state, since there will be no "opt out" clause as in the cases of the United Kingdom and Denmark. This means that two or more years after EU enlargement there could already be a far greater number of countries participating in the European Monetary Union (EMU). In the period of transition, the convergence criteria require a country to have participated in the new Exchange Rate Mechanism (ERM II) for at least two years without severe tensions.

This paper analyses the duration of the change in the central parities for the old Exchange Rate Mechanism (ERM) of the European Monetary System (EMS) in order to draw lessons on the survival of a given regime to be used for the new member states in the design of their exchange rate strategy and macroeconomic policy.

The ERM was the most prominent example of a target zone exchange-rate system. There exist an extensive literature that builds on the seminal paper by Krugman (1991) and studies the behaviour of exchange rates in target zones. The main result of the target zone model is that, with perfect credibility, the zone exerts a stabilising effect (the so-called "honeymoon" effect), reducing the exchange rate sensitivity to a given change in fundamentals. Nevertheless, in a target zone with credibility problems, expectations of future interventions tend to destabilise the exchange rate, making it less stable than the underlying fundamentals (Bertola and Caballero, 1992). Therefore, credibility (i.e., the degree of confidence that the economic agents assign to the announcements made by policymakers) becomes a key variable. In a context of an exchange rate target zone, like the EMS, credibility refers to the perception of economic agents with respect to the commitment to maintain the exchange rate around a central parity. Therefore, the possibility for the official authorities to change the central parity could be anticipated by the economic agents, triggering expectations of future changes in the exchange rate that can act as a destabilising element of the system.

In the literature, there have been several tests of credibility based on financial instruments that reflect these market expectations [see Svensson (1992) for a survey]. Svensson (1991), e. g., proposed an arbitrage-based test of exchange rate credibility. This test uses the interest rate differential between the domestic and foreign currency to show that whenever the forward exchange rate lies outside the band, one can reject the null hypothesis that the existing exchange rate is credible. On the other hand, Campa and Chang (1996) use option prices to derive two more arbitrage tests of credibility of an exchange rate. Furthermore, Fernández-Rodríguez *et al.* (2002a) develop a test for target-zone credibility that makes use of nonlinear forecastable dependencies in time series. Finally, Ledesma-Rodríguez *et al.* (2001) study the degree of credibility of the EMS using a battery of credibility measures.

On the other hand, another important line of research has aimed to estimate the probability of realignment by means of econometric techniques. To that end, some explanatory variables of that probability are considered, usually the interest rate differential, the inflation differential, the current account balance, and the unemployment rate (see, e. g., Edin and Vredin, 1993).

We depart from the previous papers by utilising duration analysis to examine the survival of exchange rate regimes in the EMR. We have applied this approach to eight currencies participating in the ERM, using weekly data of exchange rates *vis-à-vis* the Deustchmark for the 13 March 1978-31 December 1998 period, covering in the complete EMS history.

The paper is organised as follows. In Section 2 we present the data. Section 3 briefly describes the test procedure, while Section 4 presents the empirical results. Section 5 assesses the possibility of heterogeneity in the sample. Finally, some concluding remarks are provided in Section 6.

2. Data

Concern about excessive exchange rate volatility during the 1970s and its possible adverse effects on the process of European integration prompted the establishment of the EMS in March 1979. A main element of the EMS was the ERM, an adjustable peg system in which each currency had a central rate expressed in the European Currency Unit (ECU), predecessor of the euro. These central rates determined a grid of bilateral central rates *vis-à-vis* all other participating currencies, and defined a band around these central rates within which the exchange rates could fluctuate freely. In order to keep these bilateral rates within the margins, the participating countries were obliged to intervene in the foreign exchange market if a currency approached the limits of its band. For this purpose, special credit facilities were established. If they decided by mutual agreement that a particular parity could not be defended, realignments of the central rates were permitted.

It is common to divide the experience of the ERM in three subperiods [see, e. g., Higgins (1993)]. The first subperiod extends from the inception of the ERM in March 1979 to January 1987, characterized by frequent realignments to correct for divergence in economic fundamentals of the participating nations. The second subperiod (the so-called "new ERM) lasted from 1987 to the end of 1991 and coincided with increasing confidence in the ERM, the removal of capital controls, and greater convergence in the economic fundamentals. The third subperiod covers successive crises of September 1992 and August 1993, being the German unification and the recession in Europe widely accepted as the underlying causes of such crises [see, e. g., Commission of the European Communities (1993)]. We can consider a new subperiod initiated with the broadening of the fluctuation bands to ±15% in August 1993 and characterized by volatility levels comparable to those prevailing before the crisis [see, e. g., Sosvilla-Rivero *et al.* (1999)].

Table 1 shows the main realignments and changes in the ERM during the 1979-1998 period. As can be seen, the fluctuation band was originally set at ±2.25%, but a ±6% band was set for Italy and the newcomers (Spain, the UK and Portugal). After almost a year of unprecedented turmoil in the history of the EMS, the fluctuation bands of the ERM were broadened to ±15% in August 1993 (except for the Dutch guilder and the Deutschmark, which remained with the narrow bands of ±2.25%). On the other hand, as shown in Table 1, there were nineteen realignments in the EMS history, being twelve of them prior to the currency turmoil of 1992/93.

[Table1, here]

In our study we use weekly data of eight currencies participating in the ERM of the EMS: the Belgian franc (BFR), the Danish crown (DKR), the Portuguese escudo (ESC), the French franc (FF), the Dutch guilder (HFL), the Irish pound (IRL), the Italian lira (LIT) and the Spanish peseta (PTA). Given the central role of Germany in the European Union (see Bajo-Rubio *et al*., 2001), our exchange rates are expressed *vis-à-vis* the Deutschmark. The sample period runs from the 13 March 1978 to 31 December 1998 (1034 observations), covering in the complete EMS history.

From this data set we define a dummy variable called *failure*, taking value one if a regime change occurs and zero otherwise. To that end, we shall consider as a regime change each realignment, modification of fluctuation bands or entrance in the ERM. From this variable, we generate the *duration* variable, representing the time passed between two consecutive regime changes. These variables, *duration* and *failure*, define the survival-time data associated with each regime. We have multiple failure-time data or multivariate survival data. These data arise from time-to-occurrence studies when either of two or more events (failures) occurs for the same subject (currency), or from identical events occurring to related subjects such as family members or classmates.[1] In our data, the IRL has ten failures, while the ESC also changes five times. The simplest way of analysing multiple failure data is to examine time to first event, ignoring additional failures. This approach, however, is not adequate because it wastes possibly relevant information. Alternative methods have been developed that make use of all available data while accounting for the lack of independence of the failure times. Two approaches to modelling these data have gained popularity over the last few years. In the first approach (the frailty model method), the association between failure times is explicitly modelled as a random-effect term, called the frailty. Frailties are unobserved effects shared by all members of the cluster. These unmeasured effects are assumed to follow a known statistical distribution, often the gamma distribution, with mean equal to one and unknown variance. In the second approach (the so-called "variance-corrected" models), instead of including the dependencies between failure times, the covariance matrix of the estimators is adjusted to account for the additional correlation. In this paper we follow this second approach. Finally, duration is right censored when the currencies do not experience a change at the time of leaving the sample. This is due to the announcement of the fixed bilateral conversion rates in May 1998, in addition with a clear commitment to intervene in unlimited amounts to make sure that the market rates would be driven towards those conversion rates. As a result, speculation became stabilising and the market rates converged closer and closer to the fixed conversion rates, although the world was hit by major crises during the second half of 1998 (see De Grauwe *et al.*, 1999).

The summary statistics of *duration* and *failure* variables are presented in Table 2, where both are expressed in weeks. As can be seen, for all the currencies considered in this paper we have 173 observations, with a duration mean of 94.42 weeks, being the minimum duration of 9 weeks and the maximum 296 weeks. The probability of failure is 41.04%.

[Table 2, here]

Figure 1 shows the duration of the ERM regimes during the 1979-1998 period. As shown, there is a high percentage of short durations (less than 100 weeks) representing more than the 62%, while long durations (greater than 200 weeks) only account for 10.98.

[Figure 1, here]

[1]This data are frequently encountered in biomedical and other investigations. In these studies, failure times are correlated within cluster (subject or group), violating the independence of failure times assumption required in traditional survival analysis.

3. Econometric methodology

In this section, we make a short description of the basic concepts and functions used in the duration analysis. This analysis has broadly used in Labour Economics, being the main application the empirical analysis of unemployment duration [see Kiefer (1988) for a survey].

3.1. Nonparametric analysis

In the nonparametric or empirical analysis, we make use of the information contained in the duration variable, measuring the time under some state, in our case the time passed between two consecutive ERM regime changes.

Let *T*, a discrete random variable, be a spell length between two consecutive ERM regime changes. We shall have a series of data $t_1, t_2, ..., t_n$ corresponding with all the regimen changes in the ERM. The probability distribution of duration can be specified by the distribution function

$$F(t) = Pr(T<t) \tag{1}$$

which indicates the probability that the random variable *T* is less than some value *t*. The corresponding probability is:

$$p(t) = Pr(T=t) \tag{2}$$

However, in order to study duration data it is useful to define the *survival function*

$$S(t) = Pr(T \geq t) = 1 - F(t) \tag{3}$$

giving the probability of a duration being greater than or equal to *t*. On the other hand, we can also consider the *hazard function*, defined as follows

$$h(t) = Pr(T=t/T \geq t) \tag{4}$$

as can be seen, the hazard function gives the probability of a regimen change, given a survival up to *t*.

The relation of the survival function to the hazard function is given by the following expression:

$$S(t) = \prod_{s=1|t} (1 - h(s)) \tag{5}$$

The nonparametric analysis is used to estimate the unconditional hazard function that takes into account all observations with a regime change (i. e., the relative frequency of observations with $T=t$). To that end, the Kaplan-Meier product-limit method is usually employed (see Kaplan and Meier, 1958). The hazard function is computed as follows:

$$\hat{h}(t) = \frac{d_t}{n_t} \tag{6}$$

where d_t denotes the number of changes registered in time t and n_t is the population alive at the time t before the change takes place. The Nelson-Aalen estimator of the cumulative hazard rate function is due to Nelson (1972) and Aalen (1978), and is defined up to the largest observed times as:

$$\hat{H}(s) = \sum_{s=1}^{t} \hat{h}(s) \tag{7}$$

Finally, the nonparametric maximum-likelihood estimate of the survival function is:

$$\hat{S}(t) = \prod_{j|t_j \leq t} \left(\frac{n_j - d_j}{n_j}\right) \tag{8}$$

3.2. Parametric analysis

Up to this point we have been concerned with a homogeneous population, where the lifetimes of all units are governed by the same hazard function *h(t).* We now introduce the presence of a vector of covariates or explanatory variables that may affect survival time and consider the general problem of modelling these effects. Two frequently used models for adjusting hazard functions for the effects of covariates are the multiplicative or proportional hazard rate (PH) model and the accelerated failure-time (AFT) model. First, we shall consider the proportional hazard (PH) model. This model assumes that the concomitant covariates have a multiplicative effect on the hazard function:

$$h(t) = h_o(t)g(X) \tag{7}$$

where $h_0(t)$ is the baseline hazard function and $g(X)$ is the relative risk, a proportionate increase or reduction in risk, associated with the set of characteristics X. Note that the increase or reduction in risk is the same at all durations t. Also the PH model separates clearly the effect of time from the effect of the covariates and those enter rescaling the conditional probability to leave the current regime, rather than the duration of the regime. Following the usual practice in the literature, we shall assume that $g(X)= exp(\beta'X)$. Different kinds of proportional hazard models may be obtained by making different assumptions about the baseline hazard function. For example the baseline risk may either leave unspecified yielding the Cox proportional hazard model (see Cox, 1972):

$$h(t) = h_o(t)exp(\beta'X) \tag{8}$$

or take a specific parametric form: Weibull and exponential. In the former $h_o(t)=pt^{p-1}$, and in the latter h_o=1, where p is an ancillary parameter to be estimated from the data. Note that when p>1, the risk decreases over time. Alternatively, when p<1 the risk increases over time.

Secondly, we shall consider the accelerated failure-time (AFT) model. This model changes the time scale by a factor of $exp(-X'\beta)$. In this model, the natural logarithm of the survival time $ln(t)$ is expressed as a linear function of the covariates yielding the linear model:

$$ln(t) = X'\beta + \epsilon \tag{11}$$

where ϵ is a suitable error term with density $f(\cdot)$. Different kinds of parametric models are obtained by assuming different distributions for the error term. On the one hand, if we let $f(\cdot)$ be the normal density, the lognormal regression model is obtained and the assocaited hazard function is then given by the following expression:

$$h(t) = \frac{\frac{1}{t\sigma\sqrt{2\pi}}\exp[\frac{-1}{2\sigma^2}(\ln(t)-XB)^2]}{1-\phi(\frac{\ln(t)-XB}{\sigma})} \tag{9}$$

where $\phi(\cdot)$ is the standard normal cumulative distribution and σ is an ancillary parameter to be estimated from the data. On the other hand, when *f(·)* follows a three-parameter gamma density, the generalized gamma regression is obtained. The hazard function associated with this model depends on the ancillary parameters κ and σ, being extremely flexible allowing for a large number of possible shapes. In particular, if κ=0, we have the lognormal model, and if k=1we obtain the Weibull model. Finally, if σ=1 and κ=1, we have the exponential model. Therefore, the generalized gamma model is commonly used for evaluating and selecting an appropriate parametric model from the data.

Table 3 shows the parametrisation and ancillary parameters for the survival distributions defined for both the PH and AFT models.

[Table 3, here]

Each of those alternative models implies a given temporal structure in the data. For example, the Weibull distribution is suitable for modelling data with monotone hazard rates that either increase or decrease exponentially with time, while the exponential distribution is suitable for modelling data with constant hazard. In contrast, the lognormal distribution is indicated for data exhibiting nonmonotonic hazard rates, specially initially increasing and then decreasing. Finally, the hazard function of the generalized gamma distribution is extremely flexible, allowing for a large number of shapes, including an U-shape.

4. Empirical results

In this section we present the results obtained in the duration analysis of the different regimes in the history of the ERM. We first report the results from the nonparametric analysis using the Kaplan-Meier survival and hazard estimates. We then include the covariates making use of the different parametric models introduced in the previous section.

4.1 Nonparametric estimation

Table 4 reports the number of failures and the surviving population registered at each moment *t*, and the Kaplan-Meier survival and hazard functions for a small number of durations (the total number of different durations is forty-five). Figure 2 shows the estimated survivor function for all currencies. This function gives, for each duration spell, the probability of maintaining the current regime. As can be seen, this probability decreases very rapidly for the short durations (less than 46 weeks), to register then smoother variations as time increases. This behaviour suggests that for those regimens with high durations, the ERM would be relatively stable, while for the (more common) regimes associated with short durations would be more unstable. For the whole sample, the probability of maintaining a certain regime is estimated to be 0.685.

[Table 4 and Figure 2, here]

Figure 3 presents the log-log plot for the Kaplan-Meier survival function. As can be seen, this plot reveals concavity rather than linearity, suggesting that a non-monotonic hazard function could be appropriate for our data. The estimated of the hazard function in Figure 4 gives further weight to that view, since its shape indicates a positive -to-negative duration dependence and possibly lognormality.

[Figures 3 and 4, here]

4.2 Parametric estimation

Before proceeding to present the results from the parametric estimation, it is necessary to identify and measure those variables that can influence the probability of a regime change. Given the weekly frequency of our data set, the information available on fundamental variables is relatively small. In our specification we have considered two fundamental variables: the interest rate differential of each currency with respect to Germany and a credibility indicator for the exchange rate. It should be notice that these variables appeared to be an important element in the 1992-1993 ERM crisis. In addition, they are key features in the model proposed by Ozkan and Sutherland (1995) to explain an exchange rate regime switch implemented by an optimising policy-maker, since they conclude that a fully optimal outcome can only be achieved in the case where credible pre-commitments is possible and that a reduction of the interest rate differential will increase the expected survival of a fixed exchange rate.

Regarding the interest rate differential, it follows from the uncover interest rate parity that

$$E_t(s_{t+1}) - s_t = i_t - i^*_t \tag{13}$$

where s_t is the logarithm of the exchange rate (expressed as the home currency price of a unit of foreign exchange), i_t is the domestic nominal interest rate, i^*_t is the German nominal interest rate, and E_t is the expectation operator conditioned on all information available at time *t*. On the other hand, from the purchasing power parity we have that

$$E_t(s_{t+1}) - s_t = \pi_t - \pi^*_t \tag{14}$$

where π_t and π^*_t denote expected domestic and German inflation rates, respectively. Therefore, combining expressions (11) and (12) we obtain

$$i_t - i^*_t = \pi_t - \pi^*_t \tag{15}$$

Equation (15) states that the interest rate differential incorporates the anticipated differential in theα_t inflation rates between both economies. Therefore, an increase in the interest rate differential, indicating greater risk of devaluation, will generate expectations of future realignment, affecting negatively the duration of a given regime. In this paper we use three-month interbank interest rates, kindly provided by BBVA. This interest rate is usually employed in the literature to proxy expectations in the foreign exchange market.

As for the credibility indicator, based on the results in Ledesma-Rodríguez *et al.* (2001), we use a measure of marginal credibility since it renders the best indicator to proxy the perception of economic agents with respect to the commitment to maintain the exchange rate around a central parity. This credibility measure, initially proposed by Weber (1991), focuses on the ability of policy announcements to influence the public's expectations. It measures the impact of official announcements on the exchange rate and may be thought of as the weight placed on the announcement when the public forms its expectations. This credibility measure is equal to one if the policy-maker always makes fully credible announcements, and tends to zero as the announcements become non-credible. Marginal credibility () for a given currency is defined as:

$$s_t - E_{t-1}(s_t) = \gamma + \alpha_t[c_t - E_{t-1}(s_t)] + u_t \tag{16}$$

where c_t is the logarithm of the central parity, the expectation operator is conditional to the information available in t-1, and u_t is a random disturbance. Before estimating α_t, expectations on the exchange rate must be obtained. To that end, Ledesma-Rodríguez *et al.* (2001) use a Kalman filter, obtaining a different value of α_t for each time period in the sample that we shall use as our credibility indicator. Therefore, α_t is a credibility indicator computed for each currency, each week in the sample.

We have estimated the functional forms discussed in section 2 by maximum likelihood, using 173 observations and 71 changes of regimes. In the empirical estimation we have included the percentage of realignment as a further explanatory variable in order to capture the effect of the magnitude of a realignment. Table 5 contains the parameter estimates for these alternative hazard function models for the ERM. It should be notice that the estimate from an AFT model should have the opposite algebraic sign from those obtained from a PH model or from the semiparametric Cox estimation. Therefore, to compare directly the parameter estimates and to show the impact on the hazard of each covariates, we report the corrected coefficients for the parametric models. As can be seen in Table 5, all covariates are significant and have the expected sign (positive for the interest rate differential and percentage of realignment and negative for the credibility indicator) for all the alternative specifications.

[Table 5, here]

Table 5 also reports estimates of the ancillary parameters for the last three distributions. As shown, we find a significant positive duration dependence in the Weibull model, since *p* is greater than one (1.48), indicating that as time passes the probability of a realignment increases. Regarding the log-normal model, changes in σ compress or stretch the hazard function and when this parameter is large, the hazard peaks so rapidly that the function is almost indistinguishable from those like the Weibull. In our case, we obtain an estimated value of 0.928. Finally, the shape estimate (κ=0.658) allows us to test the null hypotheses κ=0 (test for the appropriateness of the lognormal) and κ=1 (test for the appropriateness of the Weibull). In Table 5 we offer the Wald test for each of this hypotheses. The results clearly suggest that lognormal is perhaps not an adequate model for these data, while there is evidence in favour of the Weibull model.

To select a parametric model we use the Akaike Information Criterion (AIC), since we cannot use the traditional likelihood-ratio or Wald test because the models are not nested. Akaike (1974) proposes penalising each log likelihood to reflect the number of parameters being estimated in a particular model and then comparing them. In our case, the AIC can be defined as:

$$AIC=-2*(log\ likelihood)+2(c+q+1) \qquad (17)$$

where c is the number of model covariates and q the number of model-specific ancillary parameters. As shown in Table 5, the Weibull model is preferred by the AIC.

On the other hand, we can use the likelihood-ratio test (LR) to select a model. All the tests have one degree of freedom. Once again, the results in Table 5 suggest that the lognormal model is not satisfactory for our data. In contrast, there is strong support for the Weibull model.

5. Heterogeneity

In order to check the robustness of our results to changes in the sample, we have analysed the possibility of heterogeneity in our sample of currencies. To that end, we have consider two potential groups with different characteristics as shown in Figure 5 and Table 6:

- a first group of currencies (that we shall denote as "core": FF, BFR, HFL and DKR), representing the 60.12% of the observations, being 103.71 weeks the mean duration and 37.5% the probability of change, and

- a second group (that we shall denominate "periphery": IRL, LIT, PTA and ESC), representing the 39.88% of the observations, being 80.42 weeks the mean duration and 46.37% the probability of change.

[Figure 5 and Table 6, here]

It is interesting to note that these two groups roughly correspond to those found in Jacquemin and Sapir (1996), applying principal component and cluster analyses to a wide set of structural and macroeconomic indicators, to form a homogeneous group of countries. Moreover, these two groups are basically the same identified by European Commission (1995) when distincting between countries whose currencies continuously participated in the ERM from its inception maintained broadly stable bilateral exchange rates against the Deutschmark, and those countries whose currencies either entered the ERM later or suspended its participation in the ERM, fluctuating in value to a great extent relative to the Deutschmark. Finally, the same two groups are found in Fernández-Rodríguez *et al.* (1999) to have relevant information helping to improve the prediction of currencies in each group based on the behaviour of the rest of the currencies, information that can be used to generate simple trading rules that outperform the moving average trading rules widely used in the markets (see Fernández-Rodríguez *et al.*, 2002b).

Figure 6 plots the estimated survival function for the currency groups. As shown, the probability of maintaining a given regime quickly decreases in the short durations (less than 40 weeks) for both groups of currencies. It is interesting to notice that for all durations the probability of maintaining the regime in the periphery is lower than in the core, with gradual changes that take place more often and are registered until the end of the period. In contrast, for the core currencies the probability of maintaining the regime is roughly constant as duration increases. This results suggest that the currencies in the core would have been more stable. Nevertheless, it should be noted that the accuracy of the estimator is better for shorter durations, since inferences about very long duration are based on fewer observations.

[Figure 6, here]

Table 7 reports the result of a log-rank test for equality of survival functions among currencies. The estimated *p*-value indicates that we cannot reject the null hypothesis of equality of survival functions. Therefore, there is not evidence of heterogeneity in the sample. As a further sensitivity test, we test for equality of survival functions among the two groups of currencies, obtaining a *p*-value of 0.06. Consequently, we can tentatively conclude that there is certain evidence of heterogeneity among group currencies in the ERM. The evidence in favour of such heterogeneity is supported by the estimation of the parametric especification including a dummy variable that takes value one if the group of currencies is periphery and zero otherwise. As can be seen in Table 8, this variable is significant and have the expected sign. In this case, the best model for our data is the log-normal model.

[Table 7 and Table 8, here]

6. Concluding remarks

In this paper we have examined the regime changes in the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS). To that end, we have applied the duration model approach to weekly data of eight currencies participating in the ERM, covering the complete EMS history. We have studied the time spells between two consecutive changes in the ERM, estimating the survival and hazard functions of such variable.

The main results are as follows. First, we have made used of the nonparametric (univariate) analysis, concluding that the probability of maintaining the current regime decreases very rapidly for the short durations (less than 46 weeks), to register then smoother variations as time increases. Therefore, for those regimens with long durations, the ERM would have been relatively stable, while for the (more common) regimes associated with short durations would have been more unstable. The probability of maintaining a certain regime is estimated to be 0.685.

Secondly, we have applied a parametric (multivariate) analysis to investigate the role of other variables in the probability of a regime change. Our results suggest that the interest rate differential with Germany and the magnitude of the realignment would have negatively affected the duration of a given regime, while credibility would have positively influenced such duration. On the other hand, we have undertaken an exhaustive analysis to compare and validate alternative models, concluding that the Weibull model would be the more appropriate to fit the data.

Third, when distinguishing between groups of currencies, we observe that those in the core are more stable than those in the periphery. A formal test for equality of survival functions among currencies indicates that there is certain evidence of heterogeneity in our sample of currencies.

We consider that our results are of interest, not only for the European experience in the 1979-1998 period, but also for the analysis of other possible target zones as the new ERM (ERM II), linking the currencies of non-euro area Member States to the euro (both current European Union Member States and future candidates) (see ECOFIN, 2000). In this sense, the importance of the euro as a main reference currency in pegged or managed floating regimes in EU accession countries has increased over the past few years. Furthermore, the Eurosystem has recently emphasised that, after joining the EU, "ERM II should not be seen as mere 'waiting room' for the adoption of the euro, but as a meaningful policy framework within which to prepare the accession economies for Monetary Union and to achieve further real and nominal convergence" (European Central Bank, 2002, p. 109). Therefore, the lessons drawn from this paper could be applied to those 12 countries currently negotiating EU accession (Bulgaria, Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Romania, Slovakia and Slovenia).

In particular, it becomes clear that the accession countries must accept the primary objective of maintaining price stability and ensure the institutional, personal and financial independence of their central banks in order to increase the credibility of their exchange-rate commitments. In addition, and giving that EU accession implies free movement of capital, to reduce their vulnerability to increasingly volatile capital flows, accession countries should push ahead with structural reforms in the financial and corporate sectors and strength supervision. Finally, in the run-up to EU accession, there should be a clear commitment to economic policy coordination and monitoring, progressing in financial consolidation without jeopardising the pursuit of disinflation.

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Table 1: Main realignments and changes in the ERM (1979-1998).	
13.3.1979	ERM starts to operate with the BFR, DKR, DM, FF, IRL, LIT and HFL. They are in the narrow band (±2.25% fluctuation), except the LIT in the wide band (±6% fluctuation).
24.9.1979	Realignment (DKR-3%, DM +2%).
30.11.1979	Realignment (DKR-5%).
23.3,1981	Realignment (LIT-6%).
5.10.1981	Realignment (DM +5.5%, FF-3%, HFL +5.5%, LIT -3%).
22.2.1982	Realignment (BFR -8.5%, DKR-3%).
14.6.1982	Realignment (DM +4.25%, FF -5.75%, HFL +4.25%, LIT -2.75%).
22.3.1983	Realignment (BFR +1.5%, DKR +2.5%, DM +5.5%, FF -2.5%, IRL -3.5%, HFL +3.5%, LIT -2.5%).
22.7.1985	Realignment (BFR +2%, DKR +2%, DM +2%, FF +2%, IRL +2%, HFL +2%, LIT -6%).
7.4.1986	Realignment (BFR +1%, DKR +1%, DM +3%, FF -3%, HFL +3%).
4.8.1986	Realignment (IRL -8%).
12.1.1987	Realignment (BFR +2%, DM +3%, HFL +3%).
19.6.1989	The PTA joins the ERM with the wide band (±6%).
8.1.1990	The LIT joins the narrow band (±2.25%). Realignment (LIT -3.6774%).
8.10.1990	The UKL joins the ERM with the wide band (±6%).
6.4.1992	The ESC joins the ERM with the wide band (±6%).
19.4.1992	Realignment (BFR +3.5%, DKR +3.5%, DM +3.5%, ESC +3.5%, FF +3.5%, IRL +3.5%, HFL +3.5%, LIT -3.5%, PTA +3.5%, UKL +3.5%).
17.9.1992	The UKL and the LIT suspend their participation in the ERM. Realignment (PTA -5%).
23.11.1992	Realignment (ESC -6%, PTA -6%).
1.2.1993	Realignment (IRL -10%).
14.5.1993	Realignment (ESC -6.5%, PTA -8%).
2.8.1993	The ERM fluctuation bands are widened to ±15%, except for the DM and the HFL.
9.1.1995	The ATS joins the ERM with the new wide band (±15%).
6.3.1995	Realignment (ESC -3.5%, PTA -7%).
14.10.1996	The FIM joins the ERM with the newwide band (±15%).
25.11.1996	The LIT re-joins the ERM with the new wide band (±15%).
16.3.1998	Realignment (IRL +3%). The DR joins the ERM with the new wide band (±15%).

Note: ATS, BFR, DKR, DM, DR, ESC, FF, FIM, HFL, IRL, LIT, PTA and UKL denote, respectivelly, the Austrial schilling, the Belgian franc, the Danish krone, the Deutschmark, the Greek drachma, thePortuguese escudo, the French franc, the Finnish markka, the Dutch guilder, the Irish pound, the Italian lira, the Spanish peseta and the Pound sterling.

Table 2. Summary statistics

	ALL CURRENCIES	
	Failure	Duration
Mean	0.41	94.42
Standard Deviation	0.49	83.32
Median	0	75
Minimum	0	9
Maximum	1	296
Observations	173	173

Table 3: Parametric survival distributions

Distribution	Metric	Survival function	Parametrisation	Ancillary parameters
Exponential	PH	$\exp\{-\lambda t\}$	$\lambda=\exp\{X'\beta\}$	
Exponential	AFT	$\exp\{-\lambda t\}$	$\lambda=\exp\{-X'\beta\}$	
Weibull	PH	$\exp\{(-\lambda t)^p\}$	$\lambda=\exp\{X'\beta\}$	p
Weibull	AFT	$\exp\{(-\lambda t)^p\}$	$\lambda=\exp\{-X'\beta\}$	p
Lognormal	AFT	$1-\Phi(\frac{\ln(t)-\mu}{\sigma})$	$\mu=X'\beta$	σ
Generalized gamma	AFT	$1-I(\kappa, \kappa \exp\{\frac{\ln(t)-\lambda}{\sigma\sqrt{\kappa}}\}]$	$\lambda=X'\beta$	κ, σ

Note: PH and AFT denote proportional hazard and accelerated failure time, respectively. $\Phi(z)$ is the standard normal cumulative distribution and *I(k,a)* is the incomplete gamma function.

Table 4. Kaplan-Meier survivor and hazard functions

Duration	Surviving Population	Failure	Net Lost	Survivor Function	Hazard Function
9	173	1	5	0.994	0.006
10	167	2	7	0.982	0.012
12	158	2	0	0.970	0.013
14	156	0	1	0.970	0
16	155	2	0	0.957	0.013
17	153	1	5	0.951	0.007
20	147	3	8	0.932	0.020
23	136	2	0	0.918	0.015
(output omitted)					
241	19	1	4	0.470	0.053
273	14	0	5	0.470	0
282	9	0	4	0.470	0
296	5	5	0	0.000	1

Table 5: Parametric estimation				
	Cox	Weibull	Lognormal	Generalized Gamma
Interest rate differential	0.265	0.165	0.136	0.155
	(3.09)	(3.15)	(3.22)	(2.87)
Marginal credibility	-0.740	-0.628	-0.809	-0.731
	(-2.20)	(-2.18)	(-3.77)	(-2.31)
Percentage of realignment	0.168	0.140	0.149	0.143
	(4.40)	(11.41)	(9.02)	(12.60)
constant		-5.521	-5.033	-5.335
		(-15.64)	(-18.57)	(-12.02)
Ancillary parameter		p=1.480	σ=0.928	κ=0.658
		(10.01)	(10.13)	(2.33)
				σ=0.764
				(5.19)
AIC	543.68	265.67	268.01	266.70
Wald test				(κ=0) 5.41
[χ (df=1)]				(κ=1) 1.47
LR[χ (df=1)]		0.93	3.28	
Observations	173			
Changes	71			

Notes: All the models, except for the Cox model, are estimation using the AFT version.
All estimated coefficients are statistically significant at the 95% level and the standard errors are adjusted for clustering on currency.
df denotes the degrees of freedom.

Table 6. Summary statistics

	CORE		PERIPHERY	
	Failure	Duration	Failure	Duration
Mean	0.38	103.71	0.46	80.42
Standard Deviation	0.49	90.52	0.50	69.40
Median	0	75	0	68
Minimum	0	9	0	9
Maximum	1	296	1	296
Observations	104	104	69	69

Table 7: Log-rank test for equality of survival functions among currencies					
Currency	Events Observed	Events Expected	Group of currency	Events Observed	Events Expected
FF	9	11.6	CORE	39	46.04
IRL	12	10.54	PERIPHERY	32	24.96
BFR	10	11.45			
LIT	10	6.13			
HFL	9	11.6			
DKR	11	11.38			
PTA	5	5.03			
ESC	5	3.26			
Total changes	71	71		71	71
L R chi2(7)	5.35		LR chi2(1)	3.41	
Pr>chi2	0.62		Pr>chi2	0.06	

Table 8: Parametric estimation with heterogeneity by group				
	Cox	Weibull	Lognormal	Generalized Gamma
Interest rate differential	0.353	0.240	0.260	0.256
	(4.10)	(6.32)	(8.97)	(8.65)
Marginal credibility	-0.823	-0.723	-0.782	-0.799
	(-2.58)	(-2.53)	(-3.62)	(-3.62)
Percentage of realignment	0.154	0.127	0.147	0.144
	(4.14)	(10.69)	(9.14)	(8.47)
Periphery	-0.616	-0.541	-0.893	-0.823
	(-1.62)	(-2.09)	(-3.63)	(-2.74)
constant		-5.366	-4.969	-5.043
		(-18.71)	(-21.28)	(-18.53)
Ancillary parameter		p=1.518	σ=0.870	κ=0.218
		(9.57)	(8.97)	(0.63)
				σ=0.832
				(5.31)
AIC	540.02	259.06	254.76	256.42
Wald test				(κ=0) 0.40
[χ (df=1)]				(κ=1) 5.14
LR[χ (df=1)]		4.64	0.34	
Observations	173			
Changes	71			

Notes: All the models, except for the Cox model, are estimation using the AFT version.
All estimated coefficients are statistically significant at the 95% level and the standard errors are adjusted for clustering on currency.
df denotes the degrees of freedom.

Figure 1: Regimen duration in the EMR, 1979-1998

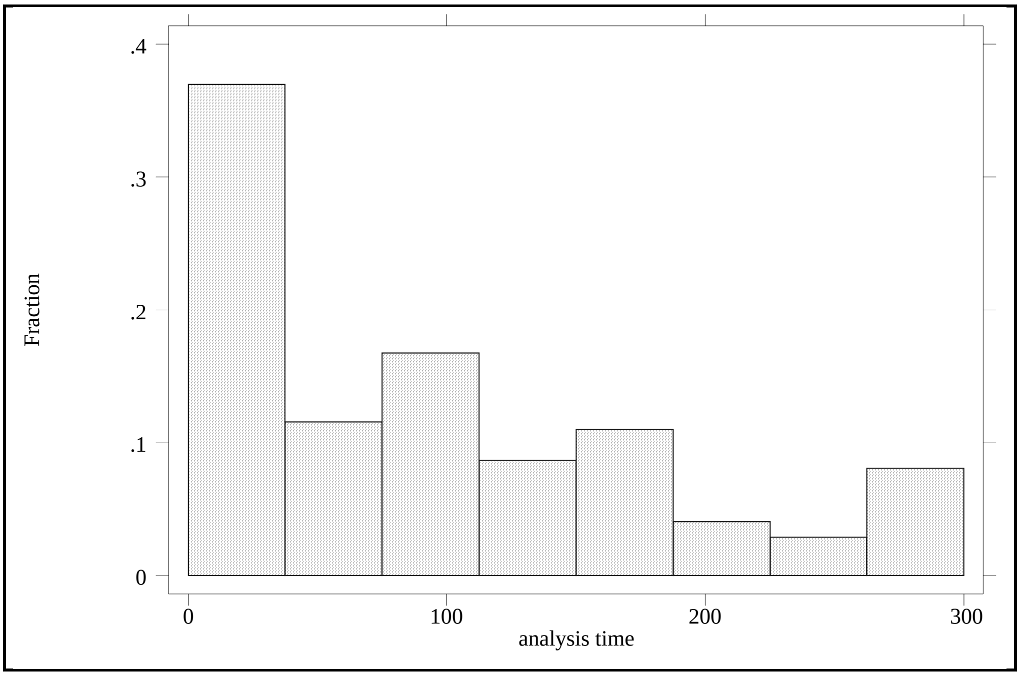

Figure 2: Nonparametric estimation of the survival function in the EMR (Kaplan-Meier estimate)

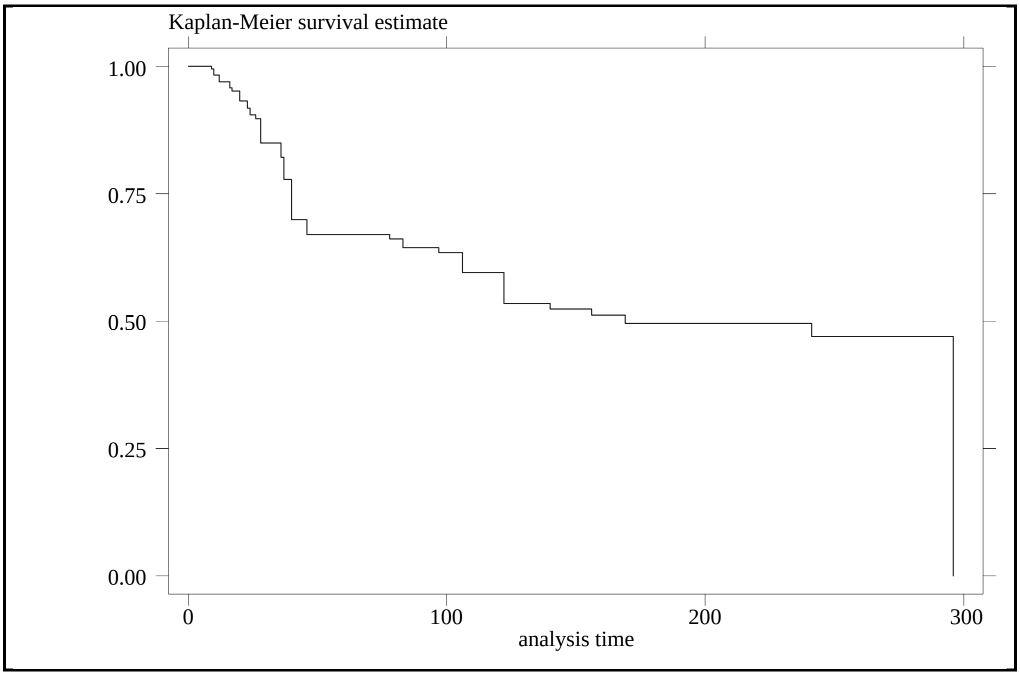

Figure 3: Log Negative Log survivor function

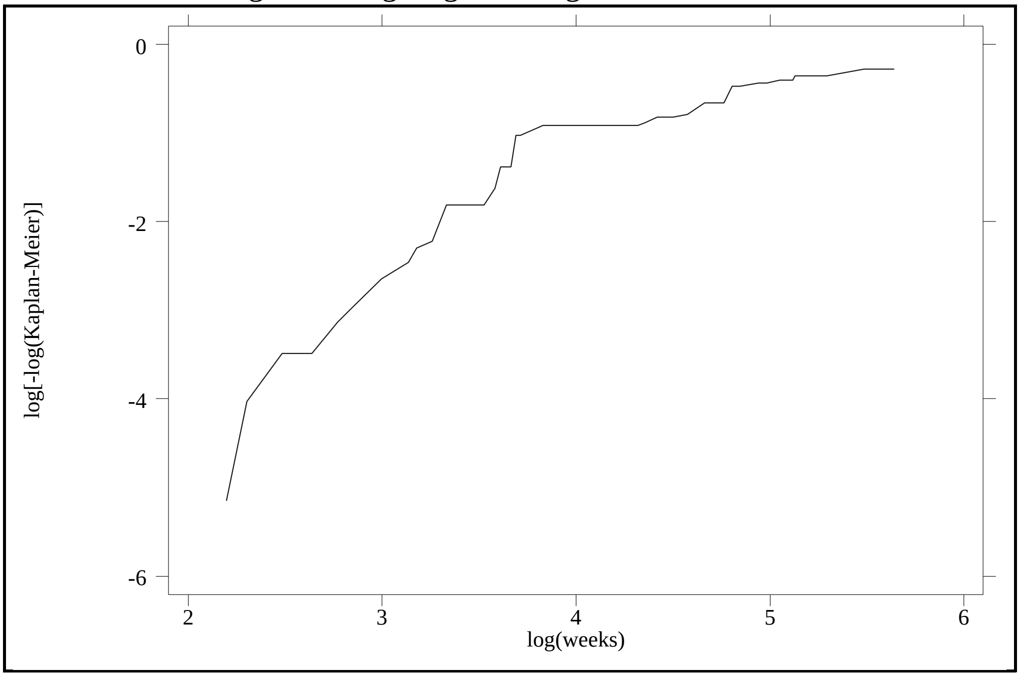

Figure 4: Non-parametric estimation of the hazard function in the ERM
(Kaplan-Meier estimate)

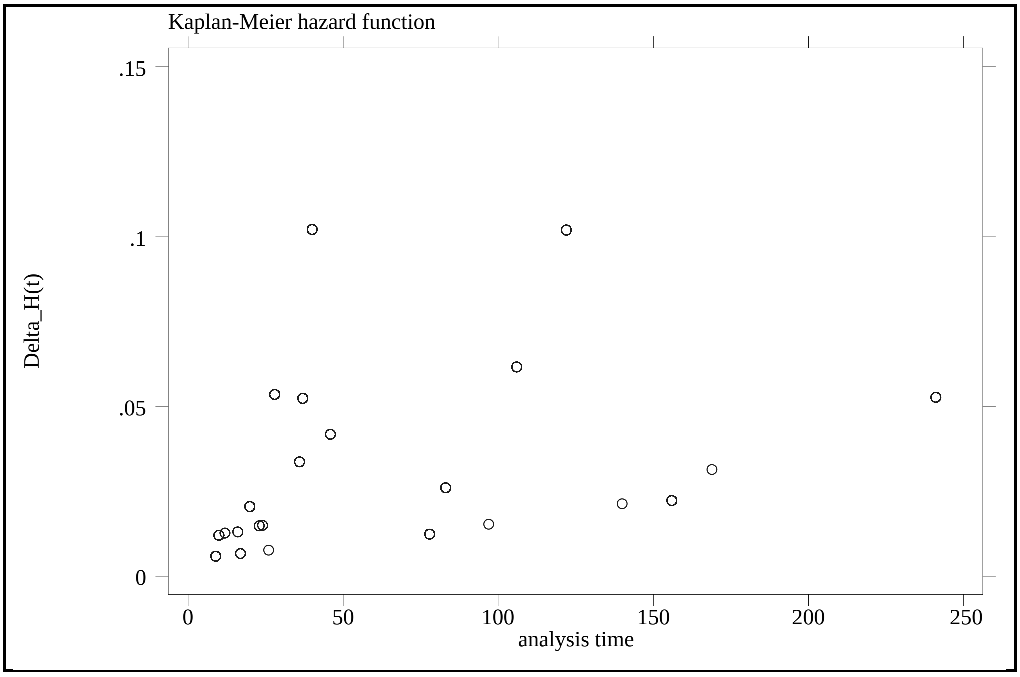

Figure 5: Regimen duration in the EMR, by groups of currencies (1979-1998)

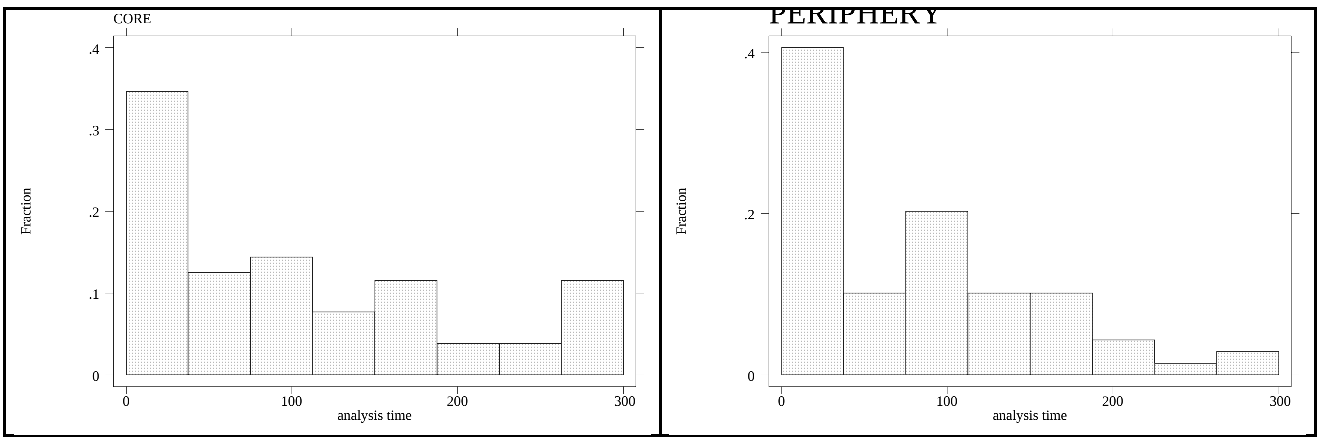

Figure 6: Nonparametric estimation of the survival function in the EMR, by groups of currencies (Kaplan-Meier estimate)