An Eclectic Approach to Currency Crises: Drawing Lessons from the EMS Experience

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ABSTRACT

This paper examines the regime changes in the European Exchange Rate Mechanism (ERM), by applying the duration model approach to quarterly data of eight currencies participating in the ERM, covering the complete European Monetary System (EMS) history. We first make use of the nonparametric (univariate) analysis, finding that the probability of maintaining the current regime decreases very rapidly for the short durations to register then smoother variations as time increases. Second, we apply a parametric (multivariate) analysis to investigate the role of other variables in the probability of a regime change. In particular we consider three alternative theoretical frameworks to select potential explanatory variables: first- and second-generation models of currency crisis and an eclectic model that combines the explanatory variables suggested by both models. Our results suggest that the Weibull specification of the eclectic model would be the more appropriate to fit our data set, finding that the real exchange rate, the interest differentials and the central parity deviation would have negatively affected the duration of a given regime, while credibility, the level of international reserves and the price level in the anchor country would have positively influenced such duration. Finally, we do not find evidence of observed heterogeneity associated to currencies with different behaviour in the sample, nor the existence in our sample of unobserved heterogeneity caused either by misspecification or omitted covariates.

JEL Codes: C41, F31, F33

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1. Introduction

The turbulence of the European Exchange Rate Mechanism (ERM hereafter) in 1992-93, the Turkish lira crisis in 1994 and 2001, the collapse of the Mexican peso during 1994-95, the Asian turmoil during 1997, the Russian currency disturbances in 1998, the crisis of the Brazilian real in 1999, and the devaluation of the Argentinian peso in 2002, have renewed the interest in the potential causes of currency crises. An extensive range of the exchange rate literature, from both the theoretical and empirical approaches, has concentrated on the modeling of exchange rate crisis.

The literature focuses around the main predictions of the canonical models, the so-called first- and second-generation models [see Kaminsky *et al* (1998), Flood and Marion, (1999) and Jeanne (2000) for recent surveys]. On the one hand, the *first generation models* stress the role of weak economic fundamentals, such as monetary and/or fiscal imbalances, in explaining currency crises. This approach is based on Krugman(1979)´s seminal paper where , under a fixed exchange rate, domestic credit expansion in excess of money demand growth leads to a gradual but persistent loss of international reserves and, given the investors' expectations, to a speculative attack on the currency. This attack forces authorities to abandon the parity because reserves are totally exhausted. The process ends with an attack because the model assumes that investors are forward-looking and, consequently, in the absence of an attack they would incur in a capital loss on their holdings of domestic money.

Different papers have extended the basic Krugman´s model in several directions. For example, Flood and Garber (1984) introduced the notion of "shadow exchange rate", namely the floating exchange rate that would prevail when reserves have fallen to the minimum level and the exchange rate is allowed to float freely; using this concept they could derive an analytical expression for the collapse time (i.e. the exact moment in which reserves are totally depleted and the government is forced to abandon the fixed exchange rate). Other models have introduced market imperfections or have relaxed the assumptions of investor's perfect foresight.

Some empirical studies have applied the first generation models to the analysis of currency crises in developing countries [see Blanco and Garber (1986), Edin and Vredin (1993), Goldstein (1996), among others], especially in Latin-American countries. The indicators used to empirically test these first generation models include international reserves, inflation, a production variable (Gross Domestic Product or Industrial Production Index), the real exchange rate and in the case of adjustable exchange-rate pegs, and the central parity around which the exchange rate can freely float.

On the other hand, some recent models points out that crises may arise without any noticeable change in economic fundamentals. Two crucial assumptions in these models are the introduction of nonlinearities and the reaction of government policies to changes in private behavior. The economic agents take this relationship into account in forming their expectations, but simultaneously their actions affect some variables to which the government policies respond. This circularity and the existence of nonlinearities give raise to the existence of multiple equilibria, some of which can be stable, others unstable, and the economy can move from one to another without any change in the fundamentals. This is the basic framework of the *second generation models* or *"endogenous policy" models*. Some relevant papers in this approach are Eichengreen and Wyplosz (1993), Obstfeld (1994, 1996), and Sachs *et al.* (1996). The empirical test of these second generation models have been applied to currency crises in industrial countries, specially in Europe, [see, e. g., Eichengreen *et al.* (1995, 1996)], using a wide set of indicators to explain these episodes such as the interest rate differentials, international reserves and stock indexes.

Finally, another line of research has addressed the question of the duration of a given exchange-rate regime. For example, Klein and Marion (1997), conduct a theoretical and empirical investigation into the duration of the exchange-rate pegs for sixteen Latin-American countries and Jamaica during a forty years period, whereas Flood and Marion (1995) extended the analysis to seventeen Latin-American countries in the same period.

In this paper, we aim to combine these two lines of research (currency crisis models and duration analysis) to assess the economic significance of the determinants of currency collapses. To that end, we depart from the previous papers by using duration analysis to examine the survival of the central parities in the ERM. We have applied this approach to eight currencies participating in the ERM, using quarterly data of exchange rates *vis-á-vis* the Deustchemark for the first quarter 1979 to the fourth quarter 1998 period, covering the complete history of the European Monetary System (EMS hereafter).

The analysis of the duration of the exchange rate regimes in the EMS is a very interesting question, given the central role of *credibility* (i.e., the degree of confidence that the economic agents assign to the announcements made by the policy makers) in a context of an exchange rate target-zone, like the EMS. If we find that the dependence on duration is positive (i.e. as time passes the probability of a change in the regime takes place increases), then we could think that economic fundamentals have played a key role in stabilizing the exchange-rate regime. This view could support the need for the strict requirements imposed by the Maastricht Treaty to the potential candidates to join the Economic and Monetary Union (EMU). Otherwise, if we find that the dependence on duration is negative (i.e. as time passes the probability of a regime change decreases), then the most important question to be considered by the authorities in determining the exchange-rate regime might have been credibility.

It must be stressed that after 1999, the EU member states which are not participating in the single monetary policy (Denmark, Sweden and United Kingdom) are being given the opportunity to prepare themselves for full integration into the EMU by linking their currencies to the euro in the context of a new, modified exchange-rate mechanism (known as "ERM II" for short). In addition, the twelve accession countries are expected to demonstrate progress towards achieving the conditions necessary to adopt the euro, including participation in the ERM II. Therefore, we consider that our analysis is 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 ERM II or other attempts to maintain regional currencies pegged.

The paper proceeds as follows. In Section 2 we review the main developments in the ERM and present the survival data. Section 3 briefly describes the methodology of duration model approach, while in Section 4 we report the empirical results. Section 5 assesses the possibility of heterogeneity in the sample. Finally, some concluding remarks are provided in Section 6.

2. Duration of regimen changes in the ERM

The EMS was created in March 1979 in a moment characterized by the excessive exchange rate volatility during the 1970s and its possible adverse effects on the European integration process. 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 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 distinguish four different subperiods in the experience of the ERM (see, e. g. De Grauwe, 2000). The first subperiod extended from the ERM inception, in March 1979, to January 1987. During this subperiod, the relatively large fluctuations bands in the EMS (compared to those in the Bretton Woods system), together with relatively small and frequent realignments, helped to reduce the size of speculative capital movements and stabilised the system. The second subperiod, the so-called “New ERM”, lasted from 1987 to the end of 1991, coinciding with increasing confidence in the ERM, the removal of capital controls, and a greater convergence in the economic fundamentals. The third subperiod covered successive crises of September 1992 an August 1993, where the evolution of the EMS into a truly fixed exchange rate system with almost perfect capital mobility led to credibility losses in a context of policy conflict among EMS countries about how to face the severe recession experienced in 1992-93. Finally, a fourth subperiod iniciated after the crisis of 1993, when the EMS changed its nature in drastic ways: the EMS gained credibility with the enlargement of the fluctuation bands to ±15% (reducing the scope for large speculative gains) and with the fixed exchange rate commitment among potential EMU member countries. As a result, speculation became a stabilising factor and the market rates converged closer and closer to the fixed conversion rates, although the world was hit by a major crisis during the second half of 1998 (De Grauwe *et al.*, 1999).

Table 1 shows the main realignments and changes in the EMS during the 1979-1998 period. As can be seen, although the fluctuation band was originally set at ± 2.25%, Italy and the newcomers (Spain, United Kingdom and Portugal) used a wider band of fluctuation (± 6%). After almost a year of unprecedented turmoil in the history of the EMS, the fluctuation bands of the ERM were broadened in August 1993 to ± 15% except for Dutch guilder and Deutschemark, which remained with the narrow bands of ± 2.25%. On 1 January 1999 the EMS ceased to exist. On the one hand, as shown in Table 1, there were nineteen realignments in the EMS history, being twelve of them prior to the currency turmoil of the subperiod 1992-1993. On the other hand, many changes affected more than one currency, such as multiple realignments or modification of fluctuations bands.

Table 1: Main realignments and changes in the ERM (1979-1998)

13.03.1979	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).
24.09.1979	Realignment (DKR –3%, DM +2%)
30.11.1979	Realignment (DKR –5%)
23.03.1981	Realignment (LIT –6%)
5.10.1981	Realignment (DM +5.5%, FF –3%, HFL +5.5%, LIT –3%)
22.02.1982	Realignment (BFR –8.5%, DKR -3%)
14.06.1982	Realignment (DM +4.25%, FF –5.75%, HFL +4.25%, LIT –2.75%)
22.03.1983	Realignment (BFR +1.5%, DKR +2.5%, DM +5.5%, FF –2.5%, IRL –3.5%, HFL +3.5%, LIT –2.5%)
22.07.1985	Realignment (BFR +2%, DKR +2%, DM +2%, FF +2%, IRL +2%, HFL +2%, LIT –6%)
7.04.1986	Realignment (BFR +1%, DKR +1%, DM +3%, FF –3%, HFL +3%)
4.08.1986	Realignment (IRL –8%)
12.01.1987	Realignment (BFR +2%, DM +3%, HFL +3%)
19.06.1989	The PTA joins the ERM with the wide band (± 6%)
8.01.1990	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%)
6.04.1992	The ESC joins the ERM with the wide band (± 6%)
14.09.1992	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%)
17.09.1992	The UKL and the LIT suspend their participation in the ERM. Realignment (PTA –5%)
23.11.1992	Realignment (ESC -6%, PTA –6%)
1.02.1993	Realignment (IRL -10%)
14.05.1993	Realignment (ESC –6.5%, PTA –8%)
2.08.1993	The ERM fluctuation bands are widened to ± 15%, except for the DM and the HFL
9.01.1995	The ATS joins the ERM with the new wide band (± 15%)
6.03.1995	Realignment (ESC –3.5%, PTA –7%)
14.10.1996	The FIM joins the ERM with the new wide band (± 15%)
25.11.1996	The LIT re-joins the ERM with the new wide band (± 15%)
16.03.1998	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, respectively, the Austrian schilling, the Belgian franc, the Danish krone, the Deustchemark, the Greek drachma, the Portuguese escudo, the French franc, the Finnish markka, the Dutch guilder, the Irish pound, the Italian lira, the Spanish peseta and the Pound sterling.

In our study we use quarterly 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-á-vis* the Deustchemark. The sample period runs from the first quarter of 1979 to the fourth quarter of 1998, therefore covering the complete EMS history.

Based on this data, we generated a dummy variable called c*hange*, taking value one if a regime change takes place and zero otherwise. To that end, we shall consider as a regime change the entrance in the ERM, each realignment or modification of fluctuations bands[1]. Using the variable c*hange*, we build a new variable called *duration*, representing the time elapsed between two consecutive regime changes. These variables (duration and change) define the survival-time data associated with each regime. Note that the same event (change) can occur on the same currency multiple times, therefore we have multiple failure-time data or multivariate survival data.[2] In addition, we have data with censuring because there are some regimen changes that had not yet finished when the EMS ceased.

Table 2. Descriptive statistics

	ALL CURRENCIES	
	Change	*Duration*
Mean	0.416	6.617
Std. Dev.	0.494	5.976
Skewness	0.34	1.004
Kurtosis	1.12	2.932
Min	0	1
Max	1	21
N. of change	64	
Observations	154	

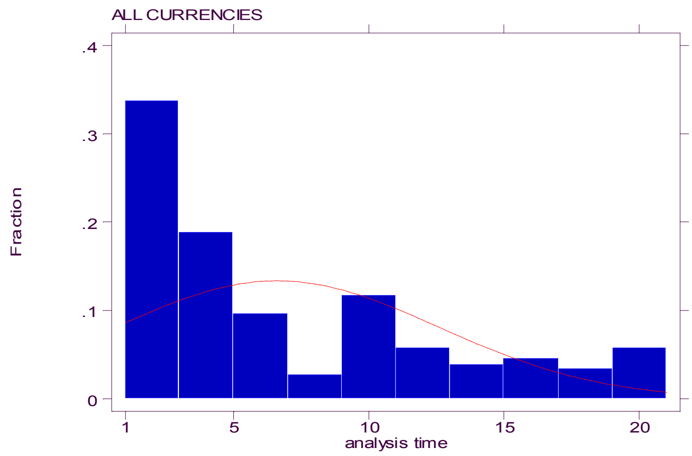
[1] In the LIT case, we also consider as change its temporary exit in the third quarter of 1992 and its re-entrance in the fourth quarter of 1996.

[2] This kind of data is 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. In our case, the 64 changes are distributed among currencies as follows: 11 for the IRL, 10 for the LIT and the DKR, 9 for the BFR, 8 for the FF and the HFL, and 4 for the PTA and the ESC.

The summary statistics for *change* and *duration* are presented in Table 2. As can be seen, for all the currencies considered, we have 154 observations. The average duration of regime changes is 6.6 quarters, being the minimum duration of 1 quarter and the maximum of 21 quarters. The average probability of change is 42%.

Figure 1 plots the duration of the ERM regimes for the whole sample period 1979-1998. As shown, there is a high percentage of short durations (less than 5 quarters), representing the 52% of the total sample, while long durations (greater than 15 quarters) only account for 9%. This result shows that the regime changes are frequent in the sample, in particular the number of changes with duration less than 5 quarters is 49.

Figure 1: Duration of regimes in the EMS, 1979-1998.



3. Econometric methodology

In this section, we offer a brief description of the main concepts and functions used in the duration models. This approach has been mainly used in Labor Economics, to study the duration of periods of employment and unemployment and the determinants of entry and exit rates [see Kiefer (1988) for a review of the literature] [3].

The duration models are used for the analysis of data which have two main characteristics: (1) the dependent variable is the waiting *time* until the occurrence of a well-defined event[4], and (2) there are predictors or explanatory variables whose effect on the waiting time we wish to assess or control.

3.1. The Hazard and Survival Functions

Let T be a non-negative random variable representing the waiting time until the occurrence of an event (change in our data). For simplicity we will adopt the terminology of survival analysis, referring to the event of interest as "death" and to the waiting time as "survival" time. We will assume for now that T is a continuous random variable with probability density function (p.d.f.) $f(t)$ and cumulative distribution function (c.d.f.) $F(t)=\Pr\{T\leq t\}$, giving the probability that the event has occurred by duration t.

It will often be convenient to work with the complement of the c.d.f, the *survival* function

[3] Duration models have been also used in the field of Industrial Organization, to analyze for example the life duration of multinational subsidiaries in the UK manufacturing industry (McCloughan and Stone, 1998), or to analyse investment ages (Licandro *et al.*, 1999). See also Sosvilla-Rivero and Maroto (2001) for a detailed study of the weekly duration of exchange rates regimes in the EMS.

$$S(t) = \Pr\{T > t\} = 1 - F(t) = \int_t^\infty f(x)dx,$$

which gives the probability of being alive at duration *t*, or more generally, the probability that the event of interest has not occurred by duration *t*.

An alternative characterization of the distribution of *T* is given by the hazard function, or instantaneous rate of occurrence of the event, defined as

$$h(t) = \lim_{dt \to 0} \frac{\Pr\{t < T \le t + dt \mid T > t\}}{dt}.$$

The conditional probability in the numerator may be written as the ratio of the joint probability that *T* is in the interval (*t*, *t*+*dt*) and *T*>*t* (which is, of course, the same as the probability that t is in the interval), to the probability of the condition *T*>*t*. The former may be written as *f(t)dt* for small *dt*, while the latter is *S(t)* by definition. Dividing by *dt* and passing to the limit gives the useful result

$$h(t) = \frac{f(t)}{S(t)},$$

which some authors give as a definition of the hazard function. In words, the rate of occurrence of the event at duration *t* equals the density of events at *t*, divided by the probability of surviving to that duration without experiencing the event.

From the above expression, we can also obtain a formula for the probability of surviving to duration t as a function of the hazard at all durations up to *t*:

$$S(t) = \exp\left\{-\int_0^t h(x)dx\right\}.$$

[4] In our case, this variable measures the time that passes between two consecutive regime changes in the ERM.

These results show the survival and hazard functions provide alternative but equivalent characterizations of the distribution of T.

One of the advantages of the hazard function is that it allows us to characterize the dependence path of duration. Formally, there exists a positive duration dependence in t^* if $dh(t)/dt>0$, in the moment $t=t^*$. This positive relation implies that the probability that a regime ends in t, given that it has reached t, depends positively on the length of the period. Thus, the longer the period, the higher the conditional probability of entering into a new regime. Similarly, there exists negative duration dependence if $dh(t)/dt<0$ in $t=t^*$. In this case, the longer the period, the lower the conditional probability of regime change.

In the above analysis we have been concerned with a homogeneous population, where the lifetimes of all individuals are governed by the same survival function $S(t)$. This analysis, which is called *"non-parametric analysis"*, is used to estimate the unconditional hazard function which registers all the observations for which there is a change, that is, the relative frequency of observations with $T=t$. For this analysis, the Kaplan-Meier estimate is widely used (Kaplan and Meier, 1958). The hazard function is calculated as follows:

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

where d_t represents the number of changes registered in moment t, and n_t is the surviving population in moment t, before the change takes place.

The Kaplan-Meier survivor function for duration t is calculated as the product of one minus the existing risk until period t:

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

3.2. Approaches to Survival Modeling

We introduce the second distinguishing characteristic of survival models- the presence of a vector of covariates or explanatory variables that may affect survival time. This analysis is called *"parametric analysis"*, and it takes into account other variables, apart from duration, that can influence the probability of a regime change. In the literature, two frequently used models for adjusting survival functions for the effects of covariates are the multiplicative or proportional hazard rate (PH) model and the accelerated failure-time (AFT) model

The first family of models –introduced by Cox (1972)- is the *Proportional Hazard model*. In this approach, the hazard function at time t for an individual with covariates x_i is assumed to be

$$h_i(t \mid x_i) = h_0(t)\exp\{x_i'\beta\},$$

where $h_0(t)$ is a baseline hazard function that describes the risk for individuals with x_i=0, who serve as a reference cell or pivot, and $\exp\{x_i'\beta\}$ is the relative risk, a proportionate increase or reduction in risk, associated with the set of characteristics x_i. This model clearly separates the effect of time from the effect of the covariates, as well as assuming that the effect of the covariates x_i is the same at all times t.

Different kinds of proportional hazard models may be obtained by making different assumptions about the baseline survival function, or equivalently, the baseline hazard function. For example if the baseline risk is constant over time, so $h_0(t)=h_0$, say, we obtain the *Exponential* regression model. Other distribution is the *Weibull distribution*, which includes the exponential as a special case. The hazard function is:

$$h(t) = \theta\lambda(\lambda t)^{\theta-1},$$

for parameters λ>0 and θ>0. If θ=1, this model reduces to the exponential and has constant risk over time. If θ>1, then the risk increases over time. Finally, if θ<1, then the risk decrease over time.

The last approach to estimate the coefficients β leave the baseline hazard $h_0(t)$ completely unspecified. This approach relies on a partial likelihood function proposed by Cox(1972).

The second family of models is the *Accelerated Life Models*. This approach is essentially a standard regression applied to the log of survival time. Using a conventional linear model, say

$$\log T_t = x_i'\beta + \varepsilon_i,$$

where ε_i is a suitable error term, with a distribution to be specified. This model specifies the distribution of log-survival for the *i*-th individual as a simple *shift* of a standard or baseline distribution represented by the error term. Different kinds of parametric models are obtained by assuming different distributions for the error term. If the ε_i are normally distributed, then we obtain a log-normal model for the T_i.

3.3. Analysis of Multiple failure-time data

The simplest way of analyzing multiple failure data is to examine time to first event, ignoring additional failures. This approach, however, is usually 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 or the failure times. Two approaches to modeling 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 modeled 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 dependencies between failure times are not included in the models. Instead, the covariance matrix of the estimators is adjusted to account for the additional correlation. In this paper we make use of these models (so-called "variance-corrected" models) in order to obtain estimation results robust to the absence of independence among observations from the same currency.

Maximum likelihood estimates of β for the Proportional Hazard Model are obtained from the partial likelihood function, $L(\beta)$, assuming independence of failure times. The estimator jruiz@cartagena.uned.es has been shown to be a consistent estimator for β and is asymptotically normal as long as the marginal models are correctly specified (Lin 1994). The resulting estimated covariance matrix obtained as the inverse of the information matrix, however,

$$I^{-1} = -\partial^2 \log L(\beta) / \partial\beta\partial\beta'$$

does not take into account the additional correlation in the data, and therefore, it is not appropriate for testing or constructing confidence intervals for multiple failure time data.

Lin and Wei (1989) proposed a modification to this naive estimate, appropriate when the model is misspecified. The resulting robust variance-covariance matrix is estimated as

$$V = I^{-1}U'UI^{-1}$$

where U is a nxp matrix of efficient score residuals. The above formula assumes that the n observations are independent. When observations are not independent, but can be divided into m independent groups (G_1, G_2, …, G_m), then the robust covariance matrix takes the form

$$V = I^{-1}G'GI^{-1}$$

where G is a mxp matrix of the group efficient score residuals.

4. Empirical Results

In this section we report the results obtained in the analysis of the different regimes in the history of the ERM. We first present the results from the nonparametric analysis using the Kaplan-Meier survival and hazard estimates. We then make use of the different parametric models introduced in the previous section in order to explore the role of different variables in influencing the probability of a regime change.

4.1 Non-parametric analysis

The estimate for Kaplan-Meier survival function is shown in Table 3 and Figure 2. For each duration, this function gives the probability of maintaining the current regime. As can be seen, these probability decreases very rapidly for the short durations (less than 4 quarters), to register then smoother variations as time increases. This behavior suggests that for those regimes with high duration, the ERM would have been relatively stable, while for the (more common) regimes associated with short durations the ERM would have been more unstable. For the whole sample, the probability of maintaining a given regime is estimated to be 0.59.

Table 3. Kaplan-Meier survivor and hazard function

Duration	*Beg. Total*	*Change*	*Net Lost*	*Survivor Function*	*Hazard Function*
1	154	11	22	0.929	0.071
2	121	15	4	0.814	0.124
3	102	18	7	0.670	0.177
4	77	4	0	0.635	0.052
5	73	0	2	0.635	0.000
6	71	1	12	0.626	0.014
7	58	0	2	0.626	0.000
8	56	1	1	0.615	0.018
9	54	4	7	0.569	0.074
10	43	7	0	0.477	0.163
12	36	1	8	0.464	0.028
13	27	1	5	0.446	0.037
15	21	0	7	0.446	0.000
18	14	1	4	0.414	0.071
21	9	0	9	0.414	0.000

Figure 2. Kaplan-Meier estimate. All currencies.

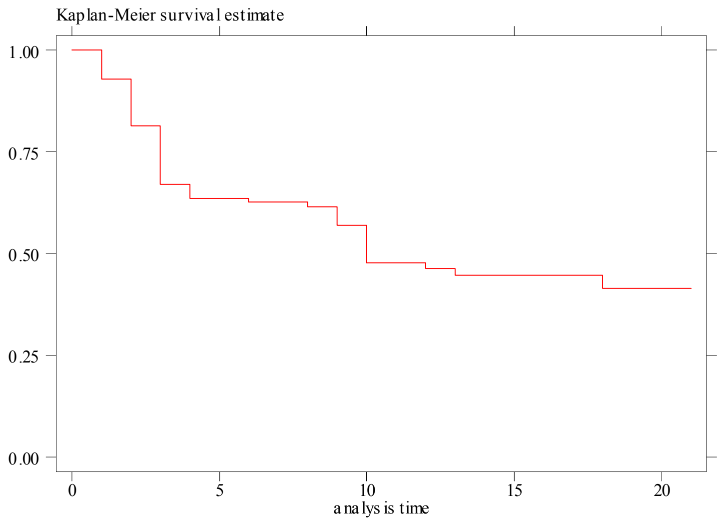


Figure 3 shows the log-log plot for the Kaplan-Meier survival function. As can be seen, this plot reveals certain linearity, at least for short durations, suggesting that a monotonic hazard function (such as a Weilbull or an Exponential function) could be appropriate for our data (Kalbfleisch and Prentice, 1980). Regarding the estimated hazard function, Figure 4 suggests a negative duration dependence, although there is evidence of positive duration around the quarters 3 and 10. Two comments are in order. First, 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. Second, the spike in quarter 10 is exclusively related to the realignment registered in 1985 due to faster Italian price increases with respect to other European countries and Italy´s large current account, so we could take this spike as an outlier. Therefore, our result suggest that a realigned exchange rate would be less durable immediately after a regimen change (as a consequence of the unstable economic environment that led to such a regime change), but once a exchange-rate regime has survived successfully for a sufficient period of time after the regime change, the probability of a regime change appears to decline.

Figure 4. Kaplan-Meier hazard estimate

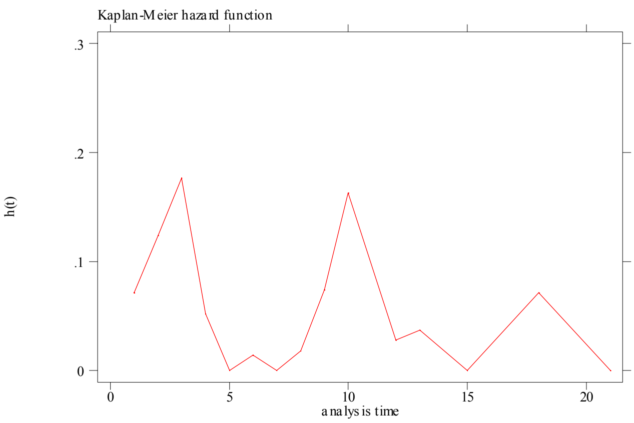
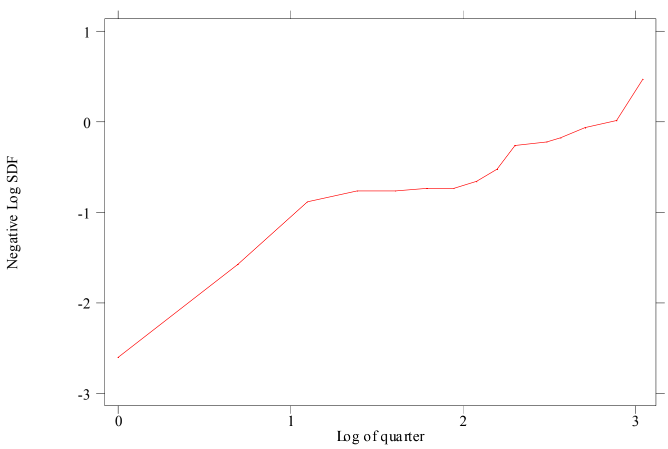


Figure 3. Log Negative Log survivor function



4.2 Parametric analysis

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. To that end, we make use of the two alternative theoretical frameworks briefly presented in the introduction (first- and second-generation models of currency crisis), as well as considering an eclectic model that combines features of both models.

Following the empirical applications of the first generation models, we start by estimating the probability of a regimen change as a function of economic fundamentals. As domestic factors, we include the money supply, the current account balance, the unemployment rate, the price level, the production level, the central parity, the level of international reserves and the real exchange rate. As for the foreign factors, we consider the money supply, the current account balance, the price level and the production level of the anchor country[5].

In contrast with the first generation models of currency crisis, second generation models emphasise the role of speculative proxies as potential causes of such crises. Following the empirical literature in this area, we examine the role of the following variables in explaining the probability of a regimen change: the level of international reserves, the interest rate differential with respect to Germany, a credibility measure, the share price index and the central parity deviation.

Finally, in an attempt to improve the explanatory power of these two approaches, the eclectic model combine the explanatory variables suggested by

[5] The exact definition of the variables as well as the data sources are detailed in the Appendix.

both models. Given that we examine the entire ERM history (from 1979 to 1998), combining features of both approaches could be a sensible option in order to take into account the possibility of different type of crises during the eighties (perhaps more related with weak county fundamentals) and the nineties (when the beliefs of foreign exchange market participants and the policy makers' reputational capital seemed to play a major role).

A class of models that has been widely used in economics and other disciplines is the proportional hazard models (see Kiefer, 1988). Therefore, we have estimated by maximum likelihood the proportional hazard specifications of the functional forms discussed in Section 3, using 154 observations and 64 changes of regime. Following a "general-to-specific" modelling methodology [see, e. g., Hendry (1995)], we started from the most general specification of hazard rate and then we simplified and re-parameterised until a parsimonious representation of the data generating process was arrived at. Table 4 to 6 contain the parameter estimates for the proportional hazard model for the ERM under the three specifications: Cox, Weibull and Exponential. Recall that a positive parameter indicates an increase in the hazard rate (that is, an increase in the probability that a given regime will end in period $t+1$, given that it lasted through period t).

In Table 4, we report the estimation results using the explanatory variables suggested by the first generation models. As can be seen, all the variables in are statistically significant at the usual level. The results suggest that an increase in the level of output (included in our specification through the industrial production index), signals stronger economic performance and then reduces the pressure on the domestic currency. Table 4 also suggests that increases in the level of international reserves significantly reduces the probability of a regimen change, while an increase in the real exchange rate (which might indicate a loss external competitiveness), would result in a higher probability of a regime change. Finally, we find that a higher price level in Germany would reduce the probability of devaluation though a reduction in inflation differentials with the anchor country [see Ötker and Pazarbaştoğlu (1997) for a similar result].

Regarding the explanatory variables suggested by the second generation models of currency crisis, Table 5 suggests that the probability of a regime change is significantly increased by an increase in interest differentials and by central parity deviation. By contrast, growing credibility appears to significantly reduce the probability of a regime change.

As for the eclectic model, Table 6 reports the estimation results. As can be seen, all the statistically significant variables that played a role in determining the probability of a regime change suggested by the previous approaches appear to influence such probability in the eclectic model, except for the level of output.

Tables 4 to 6 also report estimates of the ancillary parameters for the Weibull distribution. As shown, we find a significant positive duration dependence, since θ is greater than one (1.231 for the first generation models, 1.384 for the second generation models and 1.479 for the eclectic model), indicating that as time passes the probability of a realignment increases, in contrast with the empirical hazard function obtained using the Kaplan-Meier method (see Figure 4), perhaps due to the high percentage of short durations in our sample. The estimates suggest that the hazard rate is increasing over time at a decreasing rate (note that $1< \theta <2$) and therefore the economic fundamentals become the most important question to evaluate the stability of such a regime, supporting the relevance of the strict requirements imposed by the Maastricht treaty

Finally, in order to select the particular specification which better fit our data, we will use the Cox-residuals. We can verify the best-fitting model by calculating an empirical estimate of the cumulative hazard function, using the Cox-Snell residuals as the time variable. These residuals are defined as follows:

$$\hat{e} = -\log S(t/x)$$

where *S(t/x)* is the estimated probability of surviving to time *t*. If the fitted model is correct, these residuals, which are always positive, should have a standard censored exponential distribution with hazard ratio equal to one. We can verify this by plotting of the cumulative hazard versus the residuals and checking if the plot is a straight line with slope equal to unity and beginning at the origin. As shown in Figure 5, the Weibull specification for the eclectic model clearly satisfies the exponential requirement for most of the time, suggesting that this specification should be our preferred model.

As a further test, we have used the Akaike Information Criterion to select the best-fitting parametric model. Akaike (1974) proposes penalizing 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:

$$AIC = -2 * \log[likelihood] + 2(c + q + 1)$$

where c is the number of covariates and q the number of ancillary parameters. Although the best-fitting model is the one with the largest log likelihood, the preferred model is the one with the smallest AIC value. As shown in Tables 4 to 6, for the three parametric models, the Weibull specification is preferred by the AIC.

This criterion allows us not only to choose the most adequate functional form for the hazard rate, but also to select which of the three models (i.e. first-generation, second-generation or eclectic model) has the greater exploratory power. According to the AIC criterion, the eclectic model would be preferred. Therefore, the results suggest that the sustainability of a given exchange rate regime in the ERM was significantly affected both by fundamental variables and by investor's expectations on government behaviour.

Table 4. Parametric estimation for first generation models

	Cox	Weibull	Exponential
Ln (IPRI)	-1.911	-2.114	-2.074
	(-2.00)**	(-2.13)**	(-2.1)**
Reserves	-0.348	-0.466	-0.393
	(-5.05)**	(-5.91)**	(-3.74)**
Real ER	0.002	0.003	0.003
	(5.28)**	(6.77)**	(4.18)**
Price Index^{G}	-4.830	-5.580	-4.763
	(-2.68)**	(-3.04)**	(-3.73)**
Constant		1.221	1.036
		(1.02)	(1.04)
Theta		1.231	
		(8.07)**	
AIC	536.88	282.66	284.64
No. Observ.		154	
No. Changes		64	

Absolute z-statistics in parentheses
Standard errors adjusted for clustering on currency
* significant at 10%; ** significant at 5%
^{G} refers to Germany

Table 5. Parametric estimation for second generation models

	Cox	Weibull	Exponencial
$i\text{-}i^G$	0.255	0.336	0.271
	(3.44)**	(3.9)**	(3.92)**
Credibility	-1.217	-1.593	-1.459
	(-2.58)**	(-3.11)**	(-2.93)**
Desv. CP	0.001	0.001	0.001
	(5.5)**	(6.09)**	(5.61)**
Constant		-3.507	-2.603
		(-4.12)**	(-4.1)**
Theta		1.384	
		(9.34)**	
AIC	528.5	262.32	269.51

Absolute z-statistics in parentheses
Standard errors adjusted for clustering on currency
* significant at 10%; ** significant at 5%
^{G} refers to Germany

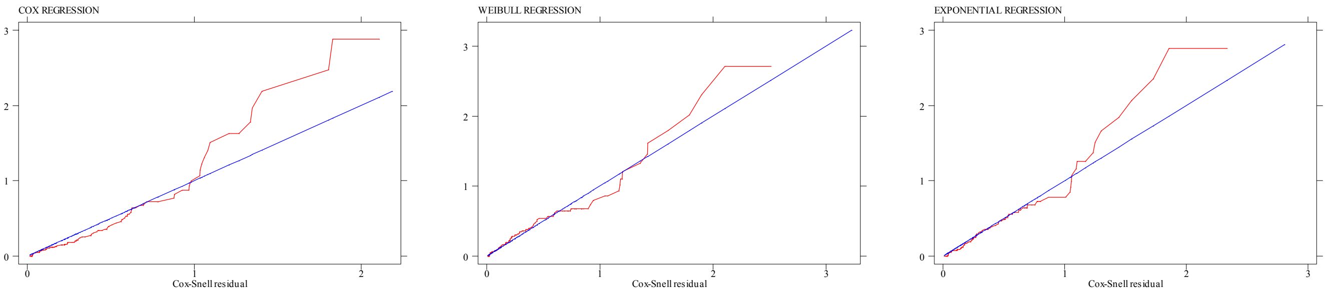
Table 6. Parametric estimation for eclectic model

	Cox	Weibull	Exponencial
Reserves	-0.400	-0.608	-0.411
	(-3.13)**	(-4.43)**	(-2.63)*
Real ER	0.010	0.014	0.010
	(3.129**	(4.4)**	(2.66)**
Price Index^G	-5.134	-5.402	-4.428
	(-2.69)**	(-2.52)**	(-2.77)**
$i\text{-}i^G$	0.187	0.257	0.201
	(2.34)**	(2.46)**	(2.55)**
Credibility	-0.992	-1.315	-1.124
	(-1.85)*	(-2.11)**	(-1.94)*
Desv CP	0.004	0.006	0.004
	(3.26)**	(4.149**	(2.84)**
Constant		0.866	1.087
		(0.42)	(0.73)
Theta		1.479	
		(9.60)**	
AIC	520.76	252.04	263.36

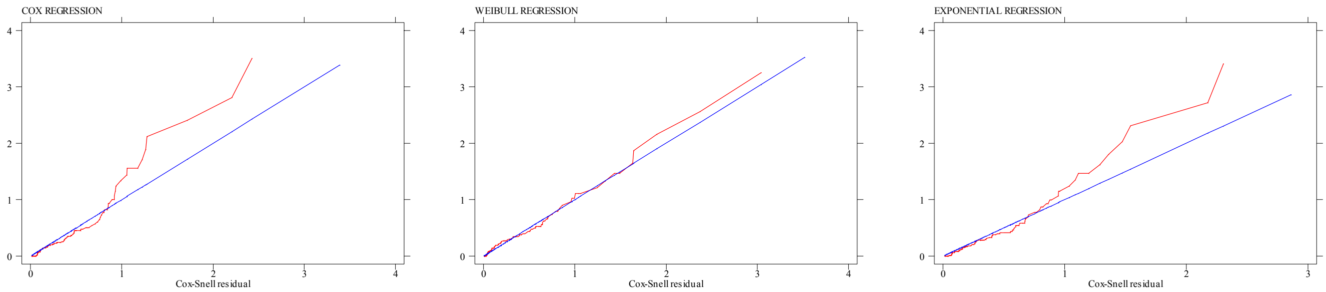
Absolute z-statistics in parentheses
Standard errors adjusted for clustering on currency
* significant at 10%; ** significant at 5%
^{G} refers to Germany

Figure 5. Cox-Snell residuals

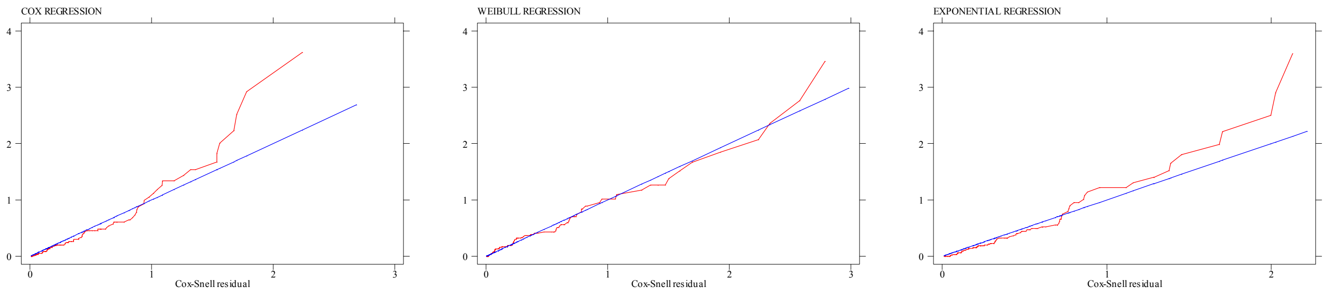
Models of first generation



Models of second generation



Eclectic models



5. Heterogeneity

In order to check the robustness of our results to changes in the sample, we have explored the possibility of heterogeneity in our data set. We can distinguish two different types of heterogeneity:

(1) observed heterogeneity that it is associated to currencies with different behavior.

(2) unobserved heterogeneity that it is caused either by misspecification or omitted covariates.

5.1. Observed heterogeneity

It is possible to identify two potential groups with different characteristics as shown in Table 7:

- A first group of currencies (that we shall denote as "core", and that include: FF, BFR, HFL and DKR), with a total of 92 observations, being 7.17 quarters the average duration and 0.38 the average probability of change.
- A second group (that we shall denote as "periphery", formed by: IRL, LIT, PTA and ESC), representing the 40.3% of the observations, being 5.79 quarters the average duration and 0.47 the average probability of change.

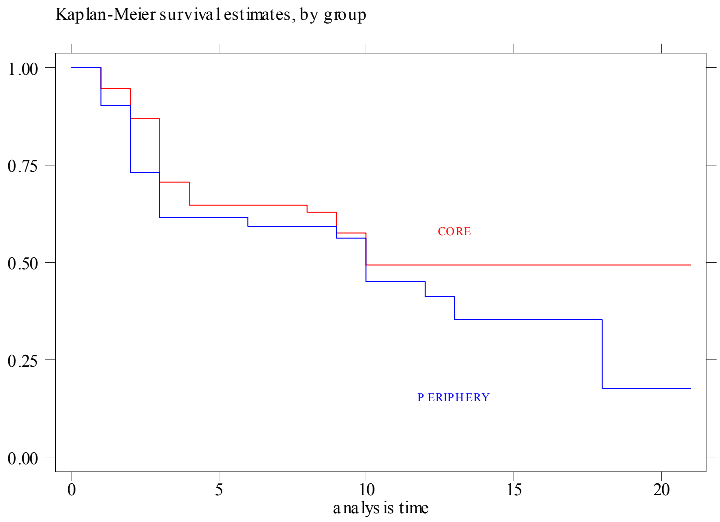
Table 7. Descriptive statistics by group

	CORE		PERIPHERY	
	Change	*Duration*	*Change*	*Duration*
Mean	0.380	7.174	0.468	5.790
Std. Dev.	0.488	6.475	0.503	5.087
Skewness	0.493	0.919	0.129	0.962
Kurtosis	1.243	2.579	1.017	3.033
Nº changes	35		29	
Observations	92		62	

It is interesting to note that these two groups roughly correspond to the distinction made by the European Commission (1995) between those countries whose currencies continuously participated in the ERM from its inception maintaining broadly stable bilateral exchange rates among themselves over the sample period, and those countries whose currencies either entered the ERM later or suspended its participation in the ERM, as well as fluctuating in value to a great extent relative to the Deutschmark. These two groups are also roughly the same found in Jacquemin and Sapir (1996), applying multivariate analysis techniques (i.e., principal components and cluster analysis) to a wide set of structural and macroeconomic indicators, to form an homogeneous group of countries. Moreover, these two groups are basically the same that those found in Fernández-Rodríguez *et al.* (1999) to have relevant information helping to improve the prediction of currencies in each group based on the behavior of the rest of 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.* (2003)].

Figure 6 plots the estimated survival functions for the currency groups. As shown, the probability of maintaining a given regime quickly decreases in the short durations (less than five quarters) for both groups. It is interesting to observe that the probability of maintaining the regime in the periphery is smaller than in the core, with gradual changes that occur more often and are registered until the end of the period. However, in the core, the probability of maintaining the regime is roughly constant as duration increases. This result would suggest that the currencies in the core would have been more stable.

Figure 6. Kaplan-Meier survival estimates by group



In order to test whether there exists heterogeneity in our sample due to different groups of currencies, we perform the Wilcoxon-Breslow test for equality of survivor function across groups [see Breslow (1970) and Gehan (1965)].

According to the results shown in Table 8, we cannot reject that equality of survival functions. Therefore, we do not find evidence of heterogeneity associated to currencies with different behaviour in the sample.

Table 8. Wilcoxon (Breslow) test for equality of survivor functions

Group	Events observed	Events expected
CORE	35	40.2
PERIPHERY	29	23.8
Total	64	64
LR chi2(1)		1.75
Pr>chi2		0.186

5.2 Unobservable heterogeneity

To address the question about the possible existence unobserved heterogeneity caused either by misspecification or omitted covariates, we take into account unobservable differences between realizations in the sample by the mean of a latent variable, called *frailty*. This latent variable could be interpreted as capturing non-economic (political, institutional, etc) idiosyncratic characteristics in the evolution of our sample of currencies that some authors have proposed to include when explaining episodes of excessive exchange rate volatility (Krugman, 1996).

Parametric specification plus covariates can only go to a certain point in explaining the variability in observed durations, being the excess unexplained variability known as *overdispersion*. A frailty model attempts to capture this overdispersion by modeling it as resulting from a latent multiplicative effect, α:

$$h(t_i | \alpha_i) = \alpha_i h(t_i)$$

where *i* refers to the *i*-th observation and *h(t)* represents the hazard function from a model we may have previously considered.

Thus unobserved differences between realizations are introduced via a multiplicative scaling factor, α_i. This is a random variable taking on positive values, with the mean normalized to one and finite variance σ^2. A crucial assumption in these models is that α is distributed independently of x and t.

Note that from the PH perspective it is very straightforward to see how α may correspond to an omitted covariate (or set of covariates):

$$h(t_i|\alpha_i) = \alpha_i h(t_i) = \alpha_i h_0(t_i)\exp\{x_i'\beta\} = h_0(t)\exp\{x_i'\beta + u_i\}$$

where $h_0(t)$ is the baseline hazard function.

The random variable u_i may be interpreted in several ways. The most common one is that it summarizes the impact of "omitted variables" or latent characteristics on the hazard rate. Alternative interpretations can be offered in terms of errors of measurement in recorded variables [for a deeper analysis on the possible interpretations of the frailty term see Hougaard (1986) and Lancaster (1990)].

In this point it is useful to make a clear distinction. Heretofore, we have taking into account differences between observations, but given that in our analysis we have considered different currencies for which different episodes are observed, then we could think in terms of a common latent effect:

$$h(t_{ij}|\alpha_j) = \alpha_j h(t_{ij})$$

for the i-th observation on the j-th currency. Taking into account this shared frailty effect would be similar to consider a between-groups random effect in a panel data model.

Estimating this shared frailty model requires an explicit assumption about the functional form of the density of α_i. Any continuous distribution supported on the positive numbers that has expectation one and finite variance σ^2 is allowed but the literature about this question usually restrict the choice between either the Gamma distribution or the Inverse-Gaussian distribution. In this paper we have selected the Gamma distribution. Once the model has been estimated we must conduct a likelihood-ratio test of the null hypothesis $H_0 : \sigma^2 = 0$. If the null hypothesis is rejected then our sample would not support the existence of this common latent effect, but if the hypothesis is not rejected then we could think that exist an unobservable random effect that captures idiosyncratic non-economic differences (i.e., political, institutional, etc.) between currencies.

In Table 9 we present the estimate of two models: (1) the reference model selected in the previous section as the best one that fitted our data; (2) the frailty model that consider the possible existence of a common latent effect between different currencies.

We can observe that the value and sign of the estimates coefficient in the frailty model are similar to those obtained in the reference model. However, the level of international reserves, the real exchange rate and the deviation central parity lose their significance when we control by shared unobserved heterogeneity. Also the Weibull distribution shape parameter θ is larger in the frailty model than in the reference model- the baseline hazard slopes upwards to a greater extent.

The σ^2 value reported in the Table 9 is the estimate of the frailty distribution variance. Note that the reference model is preferred to the frailty model according to the relevant likelihood ratio test, indicating that the frailty variance is close to zero. Hence the existence of unobserved heterogeneity in our sample is rejected.

Table 9. Parametric estimation for eclectic model with heterogeneity unobserved (Weibull distribution)

	Reference Model	Frailty Model
Reserves	-0.608	-0.611
	(-4.43)**	(-1.4)
Real ER	0.014	0.015
	(4.4)**	(1.49)
Price Index^G	-5.402	-5.504
	(-2.52)**	(-3.7)**
$i-i^G$	0.257	0.266
	(2.46)**	(4.17)**
Credibility	-1.315	-1.461
	(-2.11)**	(-2.24)**
Desv CP	0.006	0.006
	(4.149)**	(1.55)
Constant	0.866	1.057
	(0.42)	(0.78)
Theta	1.479	1.499
	(9.60)**	(9.57)**
Sigma (σ^2)		0.066
AIC	252.04	251.92
LR test[χ^2(df=1)]	0.12	

Absolute z-statistics in parentheses
Standard errors adjusted for clustering on currency
* significant at 10%; ** significant at 5%
^{G} refers to Germany

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 quarterly data of eight currencies participating in the ERM, covering the entire EMS history. In particular, we have studied the length of time that elapses between two consecutive regime changes in the ERM, estimating the survival and hazard functions of such variable.

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 4 quarters), to register then smoother variations as time increases. Therefore, for those regimes with high 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 0.56.

Second, we have applied a parametric (multivariate) analysis to investigate the role of other variables in the probability of a regime change. In particular we consider three alternative theoretical frameworks to select potential explanatory variables: first- and second-generation models of currency crisis and an eclectic model that combines the explanatory variables suggested by both models in an attempt to improve the explanatory power of these two approaches. After undertaking an exhaustive analysis to compare and validate alternative models, we conclude that the Weibull specification of the eclectic model would be the more appropriate to fit our data set. Our results suggest that the real exchange rate, the interest differentials and the central parity deviation would have negatively affected the duration of a given regime, while credibility, the level of international reserves and the price level in the anchor country would have positively influenced such duration. Therefore, the empirical evidence presented in this paper suggesting that the sustainability of a given exchange rate regime in the ERM was affected both by fundamental variables and by investor's expectations on government behaviour, might indicate that to prevent currency crises it is not sufficiently to pursue sound economic policies, but policymakers must enhance their reputational capital with respect to their commitment to maintain the exchange rate around a central parity.

Third, when distinguishing between groups of currencies, we observe that those in the core are more stable than those in the periphery. Nevertheless, we do not find evidence of observed heterogeneity associated to currencies with different behaviour in the sample. Furthermore, the existence in our sample of unobserved heterogeneity caused either by misspecification or omitted covariates is also rejected. This result strongly suggests that the ERM would have effectively acted as a true system, where common interests would have had priority over the individual ones, and only real differences (at least as perceived by market participants) could have explained the different evolution of the participant 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 linking the currencies of non-euro area Member States to euro (both current European Union Member States and future candidates, see ECOFIN, 2000), as well as for investigating other episodes of currency crisis registered in the last three decades in many countries and regions around the world.

The use of the duration analysis have allowed us to evaluate the different approaches developed in the literature of currency crises (first and second generation models, as well as an eclectic model that combine features of both models) at the same time that has been used to characterize the dependence of duration. In view of the encouraging results of the present study, some optimism about the benefits from implementing this analysis seems justified.

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APPENDIX: Definition of the variables and data sources for the parametric estimation

A) Variable names and definitions:

Dependent variable:

Probability of regimen change

Explanatory variables:

CA = current account balance (IFS, line 78ald).

Credibility = marginal credibility indicator δ_t, defined as:

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

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. Note that different value of δ_t is obtain for each time period in the sample.

E = share price index (MEI).

i = short-term interest rate (IFS, line 60c).

M = money supply: M1= local currency (IFS, line 34ª.u) + deposits (IFS, line 34.b.u)

M3= M1 + quasi-money (IFS,

ΔM = changes in money supply

M/R = the ratio money supply to reserves

Y = real income: GDP = gross domestic product (IFS line 99b.c)

IPI = index of industrial production (MEI)

P = consumer price index (IFS, line 64)

R = international reserves (IFS, line 1l.d)

UR = unemployment rate (IFS, line 67r)

B) Data sources

The data base is *the International Financial Statistics* (IFS) published by the International Monetary Fund and *Main Economic Indicators* (MEI) published by the Organisation for Economic Co-operation and Development.

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