How Homogenous are Currency Crises?

A Panel Study using Multiple-Response Models

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ABSTRACT

This paper presents formal evidence that currency episodes display heterogeneity in terms of their evolution, their impact on the inflicted economy and their links with financial, political and macroeconomic fundamentals. Limited-dependent variable models for ordered and unordered outcomes along with their heteroskedastic and random effects extensions are applied on a large panel of data comprising 40 years of monthly observations on 23 developed countries. Heterogeneity, complemented by indications of self-fulfilling expectations and noise, suggest that time and region specific predictive approaches and policy responses are more useful than trying to base analysis and policy decisions on more general patterns. Results are established with formal specification tests.

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1 Introduction

A currency crisis can be defined as an occasion of extreme speculative pressure experienced by the foreign exchange market, often, but not always, followed by an abrupt devaluation of the exchange rate. Crisis episodes, including the Mexican default in 1982, the 1992-93 ERM crisis, the 1994 Mexican crisis and Tequila effect and the South-eastern Asia crisis in 1997-98, have revived the debate about the nature, frequency and scale of the phenomenon and its impact on the broader macroeconomy. The academic interest is motivated from the problem's importance for both global investors and policymakers. Both the 'crashes' and the ensuing exchange rate policy defences can cause the collapse of a government's entire macroeconomic strategy and incur severe costs for agents managing market exposures. In an era of financial integration and globalisation, questions of both the feasibility and the timing of pre-emptive measures are of critical importance.

The purpose of this article is to assess empirically whether crises of different magnitude, geographical vicinity, process of evolution, timing and outcome, occurring in various exchange rate regimes, are induced by a common set of generating factors. This is explored using a range of Limited Dependent Variable (LDV) estimation procedures applied both universally and in various sub-samples. Subsequently, various episodes are classified according to scale and whether an attack has been successful or has been repelled. Structural differentiations among them are tested for directly.

There is no consensus among theoretical models as to the existence of a single, deterministic sequence of causally inter-related events that alter the dynamics of an economy (and the foreign-exchange market in particular) leading to speculative attacks. The survey conducted by Flood and Marion (1998) shows that most classical models approximate the idiosyncrasies of the particular wave of episodes that motivated them. The literature originated with 'First Generation Models' (FGM hereafter, see Agenor et al. (1992) for a survey). A seminal paper is Salant and Henderson (1978), who utilised the Hotelling (1931) model of exhaustible resource pricing to study attacks on a government-controlled price of gold. Krugman (1979) applied the principle to fixed exchange rates, a refinement of this model being devised subsequently by Flood and Garber (1984). FGMs postulate that the initial spark for *all* crises is the inconsistency between expansionary domestic policies and rigid exchange rate targets. More specifically, crises are instigated by a government's adherence to the *priority* of an exogenously given policy goal, as in the original Krugman (1979) model in which a steadily increasing fiscal deficit is a primary goal. Although this context is deterministic, it allows for alternating attacks and recoveries of confidence if the amount of international assets that governments are willing to commit in defence is uncertain.

Second Generation Models [SGM hereafter] are based on extensions of the Kydland-Prescott (1977) and Barro-Gordon (1983) models of time inconsistency of monetary policy; perhaps the best representation being Obstfeld (1994). These posit an indirect relationship between the occurrence of crises and fundamentals. We would still expect to see some variables assuming extraordinary values, acting as the government's motivation to apply an expansionary monetary policy and thereby jeopardising a fixed exchange rate. Nevertheless, it is reduced confidence from the public in the preservation of a fixed rate that results in difficulty to defend it and not the policy conflict itself. The circular relationship between fundamentals and expectations of inflation, which are incorporated into wages, can deteriorate fundamentals and hence exacerbate the impact of an adverse demand shock, which in turn increases the government's temptation to devalue. Therefore, since the cost of defence depends on endogenous variables, e.g. interest rates, any equilibrium is "fragile" because expectations can be self-fulfilling, and *multiple equilibria* in the exchange rate are possible. The implication is that no single process can characterise all crises and any variable (e.g. increases in short-term real interest rates, unemployment, changes in government etc) can act as a "sunspot", that is, a co-ordination device for expectations, and thereby spur a crisis if the market believes it to be pertinent. Even a phenomenally trivial event can change market sentiment and trigger an attack if viewed as the culmination of a long-lasting economic or political distress.

In the light of recent episodes and especially the 1997-98 Asian crisis, the theoretical literature incorporated a further aspect, the interaction between currency crises and financial markets. Authors like Mishkin (1992, 1996), Calvo and Mentoza (1997), and Caplin and Leahy (1994) introduced the concepts of *moral hazard*, *adverse selection* and *asymmetric information* to this literature. Financial markets' imperfections coupled with implicit or explicit bailout guarantees to banks by the state, can lead to excessive and risky lending. Exogenous shocks (such as a major bankruptcy, a recession, a stock market crash, political instability or bank panics) can then evolve into a generalised financial turmoil. Sequentially, this is transmitted to the foreign exchange market.

All these theoretical perspectives regard currency episodes as discontinuities in the foreign exchange market, inherently different from the general question of exchange rate prediction. *A priori*, it seems unlikely that rare and sharp movements occurring in crises are consistent with the actions of forward-looking speculators. Should theoretical explanations on the occurrence, extent, timing and transmission of crises be rejected by empirical analysis, this would cast doubt on the validity of the rational expectations hypothesis and the treatment of the exchange rate as an asset price. Indeed, success of the empirical literature to establish a convincing connection between crises and economic fundamentals is so far limited, reinforcing perspectives of predominance of market sentiment and self-fulfilling expectations. If this is the case, the exact timing of crises is practically unpredictable but a zone of vulnerability might still be detected. Relative severity of crises in different countries could also be forecasted by approximating susceptibility to a shock, like a global decrease of confidence. Hence, empirical findings put the validity of theoretical predictions to the test, as well as providing a tool for the prediction, management and repulse of crises.

As discussed in Section 2, identification of currency crises is a challenging task since the process appears to vary across episodes and several macroeconomic indicators are involved. This diversity, which motivated the different streams of theory, obscures the direction of causalities among variables. Therefore, a critical question of empirical prediction emerges: do crises show an adequate degree of resemblance to permit generalisations from previous experience?

To answer this, an empirical approach is needed that is capable of systematically investigating the nature of the causalities surrounding currency crises in a unified manner and quantifying the extent to which crises are similar and therefore predictable. Early empirical studies, like Blanco and Garber (1986) or Cumby and van Wijnbergen (1989), are inappropriate for this purpose since they analyse collapses of specific fixings. Those episodes are not necessarily representative of the underlying population of collapsing pegs, which again is not representative of the total population of successful and unsuccessful speculative attacks on various exchange rate regimes (for example, some pegs are abandoned without being attacked).

Closer to our aim is the "indicators" approach of Kaminsky et al. (1997), which monitors unusual digressions of a series of fundamentals and accordingly signals a crisis. However, the most rigorous efforts utilise binary Limited-Dependent Variable (LDV) models applied in multi-country panels of data; notable examples are Eichengreen et al. (1995,1996), Klein and Marion (1997) and Frankel and Rose (1996). LDV methods have the advantage that they summarise all underlying causalities in a single probability measure. While this methodology can avoid episode selection biases, the above papers (in contrast to the approach adopted in this paper) implicitly impose the assumption of homogeneity of examined episodes, without formally testing for it. The typically mediocre performance of empirical models was usually attributed to loose links with fundamentals, without allowing for the possibility of it coming, at least partly, from the inherent dissimilarity of episodes under consideration. The only counter-example is Eichengreen et al. (1995), who juxtapose revaluations with regime switches, e.g. from fixed to free float, finding the former to be mirror images of devaluations but the latter largely unpredictable. Still, they used *ad hoc* definitions of crises instead of gleaning them from the data. This practice is questionable as to the objectiveness of the selection. Equally importantly, it renders the approach inappropriate for predictive purposes since this definition can only be *a posteriori*. While the understanding of heterogeneity between different waves of crises has gained ground, the existence of structural differences among episodes of different scale and outcome, even if contemporaneous, has not to date been formally explored.

This study contributes towards this end by applying advanced econometric techniques, in the novel framework of multiple-response models, but enriched with a series of innovations that extend the most successful of previous attempts. We maintain that the failure of previous empirical studies to obtain robust findings and high out-of-sample prediction, and subsequently to establish universally applicable predictive and policy rules, is related to the degree of heterogeneity among crises. Analysis is supported by tailor-made specification tests and performance measurement. Of key importance are the data, which constitute the largest panel assembled on the study of the topic to date and are unusual in this literature in that they are of monthly rather than annual or quarterly frequency. Higher periodicity helps to achieve a more comprehensive sampling of crises, by capturing smaller and shorter duration episodes, especially the unsuccessful ones, whose effects on indices would have faded away long before quarterly or annual figures were aggregated for publishing. Thus, we examine all occasions of speculative pressure instead of just extraordinary crashes. We show that there is a trade-off between the use of updated data and the increase of noise due to the higher frequency, which seems to obscure the crisis-fundamentals relationships. As a remedy, we innovate in developing heteroskedastic and random effects extensions of the basic LDV models. Then, any sensitivity that remains has to be attributed to inherent dissimilarity of studied episodes.

Establishing that multiple equilibria exist would attest that crises are dissimilar; in that case, the list of candidate fundamentals would greatly expand, as their relevance would depend only on the market's perception. However, even advocates of this theory, like Jeanne (1997), admit that testing for multiple equilibria is limited by the prerequisite that all economic and other determinants of crises are correctly called. It is methodologically preferable to group crises on the basis of in-data criteria and test directly for structural differentiations among them. If heterogeneity is established, combined with evidence of noise, it would also lend support to the self-fulfilling expectations hypothesis.[1]

Finally, two more possibilities have to be investigated. The 'fit' of models can be compromised by instability of the determinants-crises relationship over time or cross-section. Also, temporal and geographical concentration of attacks raises the possibility that some of them may not be explicable solely on the grounds of domestic economic fundamentals. Therefore, to assert in favour of heterogeneity, structural breaks and contagion must be taken in account.

The paper is organised as follows. The econometric methodology, pre-testing analysis, and data features are presented in section 2. Sections 3, 4 and 5 contain the empirical findings along with sensitivity analysis of results. Section 3 offers a benchmark for the main findings by linking them to the findings of binomial LDV models. Section 4 presents the results of ordered models on crises of

[1] Note however that if multiple equilibria exist, the belief that an attack will alter economic policies can motivate an attack even if agents recognise the consistency of present policies with a peg. Then the incompatibility with the peg concerns *future* fundamentals *given* the attack. This notion seriously complicates the effort to detect *a priori* quantifiable indications of a forthcoming crisis.

different scale along with statistical tests. Findings of multinomial models for unordered outcomes, complementing the ordered models and also investigating the factors that distinguish successful attacks from successful defences are contained in Section 5. Section 6 concludes.

This study finds persuasive evidence of existence of structural dissimilarities amongst crises. Higher-scale episodes and major devaluations have more important links with longer-term macroeconomic fundamentals, such as real growth and growth expectations, and are facilitated by the absence of capital controls. On the other hand, contagion and inflation, if not accompanied by fundamental imbalances, are associated with more ephemeral turbulence.

2 Methodology and Data

2.1 The LDV Methodology and Extensions

Previous studies employing an LDV methodology use Binary Response Models. In those, speculative demand for foreign exchange is approximated by a continuous composite index and then transformed into a qualitative variable of just two outcomes: "tranquillity" and "crisis". However, this approach cannot account for structural differences among various crises, e.g. as grouped by unbiased in-data rules. Subsequently, the universal applicability of any revealed patterns is questionable. We propose an appropriate methodology for identifying structural differences; this involves the use of non-linear LDV models for multiple outcomes. Firstly, we ask whether episodes of different scale are governed by identical causal sequences or if differentiation among them can render gains of any sort to the estimation process. To address this question, we employ the Ordered Response Model (ORM). This novelty also addresses the critique against previous studies utilising binomial LDV models, that their results are a function of the threshold by which they empirically separate "crises" from "tranquillity". Note that in this study the threshold setting could have a stronger impact on the qualitative characteristics of the sample of "crises" because we use monthly data, which offers a realistic chance to capture smaller and short-lived crises. With this consideration in mind, our modelling strategy aims to capture as many occurrences of speculative pressure as possible, from minor repelled episodes to major crashes. Subsequently, we classify them in different "crisis" categories, according to intensity, and test for differences among them. ORM is obtained by assuming an underlying response variable y_i^*, linearly related to a vector of explanatory variables:

$$y_i^* = \alpha + \mathbf{x}_i \boldsymbol{\beta} + \varepsilon_i \qquad (1)$$

In our application, y_i^* is a "speculative pressure" index (see below). y_i^* is latent. Instead we observe a categorical variable y, here a "crisis" outcome, that is strictly ranked according to the relation:

$$y_i = m \quad \text{if } \tau_{m-1} \leq y_i^* < \tau_m \qquad m=1,\ldots,J \qquad (2)$$

where τ is a threshold. Since y_i^* is unobservable the model has to be estimated with maximum likelihood methods. The probability of an observed outcome $y=m$, given $\mathbf{x}$, is:

$$Prob\ (y_i = m \mid \mathbf{x}_i) = F\ (\tau_m - \mathbf{x}_i\ \boldsymbol{\beta}) - F\ (\tau_{m-1} - \mathbf{x}_i\ \boldsymbol{\beta}) \qquad (3)$$

The model is identified by assuming either $\tau_1 = 0$ or $\alpha = 0$, the rest of the thresholds being stochastic and estimable. The choice is arbitrary and does not affect calculation of coefficients, probabilities or significance tests. τ is of no interest here, since y^* is directly observable. Assuming a logistic or normal distribution for errors ε renders the *ordered logit* or *ordered probit* model respectively. Since the two models are equivalent and produce results comparable up to the 4th decimal, we use the probit specification throughout for reasons of statistical testing.

A further important novelty is introduced in this study. Previous studies overlooked the possibility of heteroskedasticity, even when lengthy panels were used. However, this can result in inconsistent as well as inefficient estimators. Data of higher frequency are even more likely to have heteroskedastic disturbances. Therefore, we formally test for heteroskedasticity in the estimated models with an LM test. Upon rejection of the homoskedasticity hypothesis, we estimate heteroskedastic counterparts of the univariate ordered probit models, employing Harvey's (1976) specification that allows for multiplicative heteroskedasticity. Harvey's (1976) general model specifies errors of (1) as $\varepsilon_i \sim N[0, \{\exp(\gamma' \mathbf{w}_i)\}^2]$, where $\boldsymbol{\gamma}'$ is a parameter vector and $\mathbf{w}$ the vector of all variables entering the skedastic function. This is preferred over simple heteroskedastic models since it can accommodate various forms of heteroskedasticity and it can also address a more general functional form problem.

We also estimated models for random effects in panel data to account for month-specific idiosyncrasies. Random effects test for parameter heterogeneity by treating α of (1) as a random variable. If this exists, errors are serially correlated across time and resulting estimates are consistent but inefficient. Previous studies using pooled data ignored this possibility. We used the equicorrelated model evaluated by the efficient computational algorithm of Butler and Moffitt (1982).

However, prospective benefits from the use of the ORM in comparison with its binary counterpart are limited by the Parallel Regression Assumption. More specifically, in the ORM, the coefficients $\boldsymbol{\beta}$, are constrained to be all the same across outcomes and equal with the coefficients obtained from the respective binomial model. If this assumption is violated, the ORM is unattainable and outcomes cannot be ordered with respect to $\boldsymbol{x}$, even when y is putatively ordered *a priori*, as in this application. Anderson (1980) refers to that as *multidimensionality* of the relationship between $\boldsymbol{x}$ and y, meaning that more than one linear functions are needed to describe it. This assumption is formally tested; in the case of rejection alternative models for unordered outcomes are examined.

Next, the problem of empirically defining "crisis" (y*) has to be dealt with. Criteria employed to determine episodes for study have to be in-data, so that models are not inappropriate for prediction. Also, voluntary devaluations have to be excluded. Identification of currency crises is a delicate task since the process appears to vary across episodes and several economic indicators are involved.

Symptoms accompanying different episodes include reversals of capital inflows, bankruptcies of banks and non-financial corporations, government bailouts, repudiation of international debt, excessive volatility in all capital markets and usually sharp declines in GDPs in the aftermath of crises. However, the eruption of a crisis can be epitomised to the manifestation of "speculative pressure". The low empirical performance of models of exchange rate determination though, impels to approximate "speculative pressure" by an *ad hoc* construction, which extends the Girton and Roper (1977) model. This approach endorses the spirit of theoretical models and it is also capable of capturing attacks on exchange regimes less rigid than pegs. According to it, excess demand for foreign exchange is manifested through up to three non-mutually exclusive channels, namely devaluation, sales of reserves and/or raising of interest rates. A weighted average of some or all of these serves as the 'latent' variable y_i^*. Then, crisis is defined as an observation larger than a certain multiple of a standard deviation above the in-sample mean. The rule is explained analytically in each model, as it differs among them. We follow Eichengreen et al. (1995, 1996) in referring to this as the

‘Exchange Market Pressure’ index (EMP) and in scaling the variables against those of Germany, chosen as the reference country because of its post-war monetary stability. Hence:

$$EMP_{i,t} = [\ (\ a\ \Delta\ s_{i,t}\ ) + (\beta\ \Delta\ (\ i_{i,t}\ -\ i_{R,t})) - (\gamma\ (\Delta\ r_{i,t}\ -\ \Delta\ r_{R,t}\ )\ )] \qquad (4)$$

Where, *s* is the nominal bilateral exchange rate w.r.t. the DM, *i* the short-term interest rate and r a ratio of international non-gold reserves to monetary liquidity, usually M1; all variables enter in differences of natural logarithms. Δ denotes the rate of change, subscript *R* denotes the relevant values of the reference country and α, β, γ are positive constants acting as weights.

Scaling provides some form of “standardisation” but entails the hazard of rendering the data-based episode selection endogenous to movements in the reference country. Hence, we also construct an index with no reference country used:

$$EMP_{i,t} = [(\ a\ \Delta\ s_{i,t}\ ) + (\beta\ \Delta\ i_{i,t}\ ) - (\gamma\ \Delta\ r_{i,t}\ )] \qquad (5)$$

Here, exchange rates are typically expressed against the U.S. dollar, which implicitly sets U.S. as the reference country, at least as far as the measurement of the exchange rate is concerned. However, there is no theoretical or methodological justification as to why any country provides superior monetary stability to use as a basis of comparison for other countries’ variables.[2] Furthermore, given the special weight of U.S. to the global economy, it is reasonable to assume that developments in the U.S. would, to some degree, be passed on and added to shocks occurring to any other reference country. Thus, variants of both schemes are employed.

A further innovation is applied to our EMP indices. Most authors use weights α, β, and γ for “equalising” volatilities of the 3 series so that no single component dominates the index. Although plausible, this is clearly an *ad hoc* practice and it could seriously affect results since use of different metrics may lead in capturing different “crisis” situations. We test the weighting scheme in our comprehensive sensitivity analysis. However, components and especially exchange rates and reserves also have large differences of scale, and thus of variability, across countries as well. Furthermore reserves are measured in U.S. dollars while M1 and M2 are measured in local currency, which, in

[2] In fact, the choice of Eichengreen et al. (1995, 1996) of Germany is undermined (in theory) by the event of the German unification: any idiosyncratic shocks that may prevailed as an effect of that event translate into base comparators for identifying other episodes; we test that.

addition, is a different multiple of 10 of each country's currency unit in order to agree with other macroeconomic measures. Therefore, the use of a single volatility measure for all countries' data is bound to create cross-country biases in favour of larger-scale observations as use of differenced data cannot eliminate all of the difference in scale. In order to avoid averaging and biases, we compute *country-specific means and respective standard deviations for each of the three components* of the index. Then these country-specific weighted series are integrated in a new single index as in:

$$EMP_{i,t} = [(s_{i,t}/3\sigma^{s}_{i}) + ((i_{i,t} - i_{G,t}) / 3\sigma^{int}_{i}) - ((\% \Delta r_{i,t} - \%\Delta r_{G,t}) / 3\sigma^{r}_{i})] \quad (6)$$

We can now show the temporal and geographical allocation of episodes for study gleaned by our technique. For example, the ordinal index is assembled by defining "episodes of lower size" as deviations of 1.5 to 2 standard deviations larger than the sample mean of this EMP index and "major episodes" as deviations larger than 2 standard deviations apart from the mean. The index leaves 8880 non-missing observations; after applying the 'exclusion window', around 4200 are left (missing data preclude a few to be used in estimation); 'crisis' observations are 384, of which 170 are "lower scale" and 214 are "higher scale". The large number of crisis observations reflects our strategy to examine "experienced speculative pressure" in the broad sense and not just extraordinary crashes.

Figure 1 shows the allocation of smaller and larger scale episodes per country. The picture sketched partly departs from the stereotypes of which countries are prone or immune to attacks. This is the combined effect of the multi-dimensional manifestation of speculation pressure in the index, the different number of missing values for each country, the different scale and upshot (success/repelled) of gleaned episodes and the ordinal classification. However, "safe havens" experienced mainly repelled episodes and not crashes. Temporally, "speculative pressure" exhibits some peaks, which coincide with major events known to have influenced the "mature" currency markets: the crises of the dollar in late 1960's that ultimately led to the floats in 1972 and early 1973, the first and second oil crises of the 1970's, the Latin America debt crisis of 1982, the U.S. interest rates rise and appreciation of the dollar in 1983-84 and the destabilisation of EMS in 1992 after the attacks on the British pound and the Italian lira. Industrial countries seem to some extent unaffected by the Mexican crisis in 1994 but more influenced by the Asian crisis of 1997-98.

Despite improvements, the construction of the EMP index is constrained by two important data-related limitations. Firstly, Klein and Marion (1997) note that in a world of risk-neutrality and perfect capital mobility the probability of devaluation should be given by interest rate differentials. However, several factors may hinder the correspondence of interest rate differentials and expected rate of depreciation, such as the existence of controls on capital account transactions, risk premiums and interest rates being set by authorities instead of being freely determined by a mature market to reflect market conditions. Thus, the argument for including interest rates among constituents of speculative pressure is questionable. Secondly, Eichengreen et al. (1996) remark that reserves data may not capture foreign exchange intervention ably since they omit or inadequately reflect factors like off-balance sheet transactions, third-party intervention, stand-by credits and foreign liabilities.

These two points underline the problem of subjectivism in the choice of the components of the index, as well as the definition of what constitutes "speculative pressure" per se. One could argue that the imposition of capital controls is an equally informative indication of mounting of speculative pressure and it could substitute reserves in the EMP index, as the two are alternatives for a Central Bank defending its currency. [3] Then, reserves' losses could be included among explanatory variables. This ambivalence of the direction of causality raises methodological and economic questions and it casts doubts on the correct specification of the model as a whole. Several authors have advocated the lagging of regressors by one period as a possible remedy or partial attenuation of the problem of interdependence. Clearly, this is not a theoretically founded and sound solution.

The above points highlight the difficulty in deciding whether or not an attack, especially an unsuccessful one, has occurred and capturing it with a composite index. The question of whether we

[3] However, the quantification of capital controls, so that they can be used as an ordinal measure-component of an index, is a problem without an obvious solution. The IMF has constructed some semi-continuous indices by aggregating several categories of restrictions that central banks impose on the capital account. Choice and quantification of all these elements is dubious. Also, these series are available only for the last few years so that use with a long data set is impossible. Furthermore, Klein, and Marion (1997) note that false invoicing, black market transactions and other measures taken to circumvent capital controls slowly erode the reserve position and repress the policymaker's ability to maintain a fixed parity.

are interested in attacks or real episodes of devaluations is relevant. For example, a monetary authority pays a cost when it loses reserves or raises interest rates to levels incompatible with its targets of internal policies, even when no devaluation occurs. On the other hand, an international investor is naturally interested exclusively in the present and future levels of exchange and interest rates. Authors like Frankel and Rose (1996) and Klein and Marion (1997) chose to focus exclusively on exchange rate episodes that include devaluations. We implement this specification too but we integrate it in our wider modelling strategy. We use it alongside EMP indices in order to limit the danger of including instances of voluntary abandonment. This is a prerequisite for pronouncing on the heterogeneity question. Again, we let the sample select episodes for study instead of defining them *ad hoc*. Hence, "actual" episodes are detected by the use of only the exchange rate, for example:

$$\Delta s_{it} > \kappa\, \sigma_i^{\Delta s} \textit{ and } \Delta s_{it} > \lambda \qquad (7)$$

where $\sigma_i^{\Delta s}$ is the standard deviation of Δs and κ, λ are positive constants. Frankel and Rose (1996) and Goldfajn and Valdes (1997) used variants of this criterion. Its logic is to capture instances in which the devaluation is both extraordinary, after conditioning on the inflation rate, and also large enough to noticeably reduce the purchasing power of a currency. This implies a short-run alteration of the real exchange rate *e*, providing an equivalent definition of a crisis. This specification gives rise to a whole sub-category of models focused on explaining and signalling crashes exclusively. All of the aforementioned variations have been applied to this family of models too and respective sensitivity tests have been performed. Analysis is accompanied by performance measurement; in particular, the Akaike Information Criterion (AIC) is preferred over pseudo R^2's as it is appropriate for comparing non-nested models. Lower values of the AIC suggest better fit.

In this framework, we can now propose a second important categorisation to be investigated, the one between successful and repelled attacks. If it is established that crises are related to any economic, financial or political fundamentals of the local and peripheral economy and its phase, it is plausible to ask if these fundamentals help to determine whether the Central Bank will repel an attack or a devaluation will occur. For this task, we need a model able to discriminate between events that are structurally different. A suitable choice is the *Multinomial Logit Model* (MNLM). In contrast with the ORM and other multinomial alternatives for unordered outcomes, such as the McFadden (1973)

Conditional Logit or the Discrete Choice Model derived from Luce (1959), MNLM allows its coefficients $\boldsymbol{\beta}_m$ to differ for each outcome, depicting the possibility of structural differences among the determinants of various outcomes. The basic MNLM can be presented as:

$$Pr\,(y = y_m \mid \boldsymbol{x}) = \exp(\boldsymbol{\beta}_m^{\mathrm{T}}\, \mathbf{x}_i) \Big/ \sum_{j=1}^{J} \exp(\boldsymbol{\beta}_j^{\mathrm{T}}\, \mathbf{x}_i) \quad where\ \boldsymbol{\beta}_1 = \mathbf{0} \tag{8}$$

The underlying structural equation is assumed to have i.i.d disturbances (the individual heterogeneity terms) with extreme value distribution. The constraint $\boldsymbol{\beta}_1 = \mathbf{0}$ is imposed for reasons of identification, with j=1 being "tranquillity" in this notation; as in ORM, the choice is arbitrary. Since the MNLM permits the coefficient vector to differ for each outcome, it is not constrained by the PRA. Thus, it also offers a cross checking of the results of the ORM and an alternative for imposing classification structure onto the data. Even if this classification is putatively considered strictly ranked *a priori*, ordinality need not be formally established.

Now we have to devise quasi-qualitative criteria to empirically divide "successful" and "failed" attacks. We use novel in-data criteria. "Successful" describes an attack that ends with an abrupt and sizeable devaluation of the exchange rate. Although the exchange rate is a focal point in the process, the employment of the composite indices, which are also influenced by interest rates and international reserves, minimises the risk of capturing voluntary devaluations instead of genuine instances of crisis. So, a "successful" attack is postulated to have occurred when: 1) the EMP index signals an instance of crisis, of whichever sort, and 2) at the same month, an abnormal devaluation of considerable magnitude, as captured by the a dual exchange rate criterion of the sort of (7), has occurred in the country. If the composite index signals "1" but the devaluation dummy does not, this is said to be a "failed" attack. The rule is detailed for each model separately as we have tried several definitions.

In estimating the model, we can formally test whether the two "crisis" outcomes are *indistinguishable* with respect to the variables in the model indicating that the two outcomes can be combined. We construct the LR test of indistinguishability suggested by Long (1997, p.163) as follows: firstly, the observations with outcomes "successful attack" and "failed attack" are selected. Secondly a binary logit is estimated on the new sample. Then a LR test of simultaneous insignificance of all slope coefficients is conducted in the binary logit. If the specification is significant, the presence of multidimensionality is established.

2.2 Definition of the Explanatory Variables

The choice and construction of the explanatory variables is guided by the requirements of the theoretical models, taking into consideration data availability. Unusual movement in some or all of these variables is reported to have preceded most crisis episodes. We now explain how our variables are linked with theoretical models and empirical regularities and, accordingly, what impact they are expected to have. We employ the widest range of explanatory variables to have appeared in the literature, combining the successful results of previous empirical studies with innovative definitions of variables. Regressors can be divided in two broad categories: variables mimicking the factors invoked by FGM, and fundamentals acting as sunspot variables and spurring self-fulfilling crises as described in SGM. This classification should not be considered rigorous, as several variables have been invoked in both classes of models.

1st Generation Models Regressors

Accumulated real exchange rate (RER) misalignment: an overheated economy (fiscal deficits and current account deficits often occur simultaneously) combined with a peg or a managed float will lead to real appreciations. Similarly, free-floating exchange rates may become inconsistent with the level of economic activity if adjustments lag due to market inefficiencies. Expectations-driven capital inflows can also lead to appreciations, even in the absence of a real interest rate differential. A simple RER index as a measure of appreciation is tried against the percentage cumulative deviation of the RER over the last 24 months to account for lagged mean reversion to an aligned (competitive) exchange rate, which is showed in empirical studies to endure up to more than two years. The finding of Klein and Marion (1997) that the monthly probability of abandonment is increased by time already spent on the peg, obliges our measure to account for the duration of the misalignment as well as its size. Cumulative forms are also computed as a surplus over the pre-2-year 60-month average RER, to address any cross-country significant inequalities of accumulated misalignment in the base year and thus preserve the comparability of magnitudes over time and across countries. A prior appreciation (increase in the RER index) is expected to increase the risk of an attack, i.e. the coefficient should be positive and large, since appreciation has been depicted in the literature as a summary variable.

The current account surplus or deficit (as a percentage of GDP). Most attacked countries are reported to have experienced trade deficits and capital flights pre-crisis. The variable's effect is expected to be negative but mediated by the inclusion of the RER appreciation variable. Still, it can reflect differences in the external sector policies across countries, vis-à-vis a given appreciation, as well as discrepancies in the relative price of tradeables to non-tradeables not picked up by the RER.

Growth of domestic credit (as a percentage of nominal GDP). It can be, necessitated by the need to finance a fiscal deficit or, under the light of the recent Asian crisis, to support a problematic financial sector. The importance of this factor is inflated by the unrealistic assumption of the original Krugman (1979) model that there is no access to international capital markets. In modern, open economies, as those in the sample, its impact is expected to be positive but possibly not decisive.

Budget surplus(+) or deficit(-) as a percentage of GDP: It may also be a poor indicator of crises if sufficient international credit lines are available, so that the central bank need not monetize a fiscal deficit and thus trigger an instant monetary expansion. We expect a deficit to have a negative effect.

The presence of capital controls: the complementarity of capital controls with interest rate and reserves' changes as tools for a Central Bank defending a currency leads us to expect a significant correlation among them and the variable to be an important predictor of crises. However, its sign is uncertain, as capital controls are a co-ordinating device for synthesising conflicting expectations of the market. If markets perceive the restriction of certain capital transactions as a sign of difficulty of the authorities to face capital outflows, or if controls are imposed *post-factum* to limit panic market reactions and a recurring attack, the effect of the variable should be positive. If the central bank has credibility and fundamentals are healthy enough, the measure may be deemed sufficient to fence short-term reversible outflows and speculators may turn their attention elsewhere; then, the impact of the variable should be negative. The constructed measure is necessarily imperfect since Central Banks also impose capital restrictions, which are not explicit and thus detectable in the books.

The state of the banking sector-financial crisis. Complexity of financial markets and involvement of many institutional factors makes the empirical modelling of financial turmoil difficult. We follow previous approaches and approximate it by bank credit to the private sector as a percentage of GDP. The rationale is that a rapid increase of bank credit signifies that the economy has moved into a vicious boom-and-bust cycle favouring the undertaking of excessive and risky investment. Runs on the currency are just a by-product of the collapse of the financial bubble and the abrupt reversal of massive capital inflows, often accumulated in the pre-crisis years. The phenomenon, modelled by Dooley (1997), can occur even when macroeconomic fundamentals are healthy and there is little motivation for future expansionary policies (low unemployment, high growth) and it is exacerbated if *bandwagon effects* exist. The variable is expected to have a positive impact.

Money growth: All models predict that monetary expansion, whether used to relieve pressure on financial organisations and the real economy or to serve a budget deficit, will, sooner or later, inevitably lead to an increase in the price level. This can only be balanced by a change in the exchange rate if competitiveness and reserves are to be kept constant. Last-minute corrective action to the money supply is usually ineffective, especially if sterilised. For the variable to produce an abnormal and abrupt devaluation, detectable as a crisis by our EMP indices, the rate of money creation has to be significantly in excess of the percentage of deprecation allowed by the exchange rate regime (zero in a fixed peg but significantly higher in managed or free floats). Otherwise, a high inflation rate, as a prominent and recognisable sign of erosion of competitiveness, can be matched by a similar deprecation rate, which evolves in a relatively smooth and predictable pattern, without provoking attacks. We define money creation as the percentage change of real M1 or M2 (M1/P, M2/P). We test empirically to discover whether a narrow or broader liquidity basis is more relevant.

2nd Generation Models Regressors

Growth expectations/ real growth Both a lack of growth and poor growth expectations have been depicted as a possible motivation to abandon a system of fixed exchange rates or, more generally, to adopt a more expansionist policy. Following Persaud (1998), to the extent that the capital market is efficient, growth expectations can be approximated by the 1-month change in equity prices lagged 3 months. Both forms of the variable are expected to affect negatively the likelihood of an episode.

Share Prices A nominal share prices index for each country is an alternative indirect indicator of both financial turmoil and growth expectations. A fall in the attractiveness of domestic stocks will, to a certain degree, translate into a wave of capital flights and, in a longer-term perspective, also signal weak growth prospects, thus prompting a devaluation. However, Eichengreen et al. (1995) report that on the *immediate* wake of an event, stock prices may rise to reflect the favourable impact of the forecasted devaluation on exports-oriented firms. Therefore, since the variable is very volatile and adjusts rapidly, use of concurrent values with higher frequency data could yield a positive coefficient in some instances.

Electoral victory or defeat of the government. In the context of SGM, the political commitment of the government to the exchange rate regime is among fundamentals whose perceived vulnerability can trigger an attack. Changes in office are clearly chances for speculation or even herding behaviour on this commitment. Furthermore, the turns of the political circle may cause lax monetary and fiscal policies. Thus we would expect attacks to tend to coincide with elections. On the other hand, Klein and Marion (1997) note that governments can gain credibility and reputation by keeping up a peg. Thus, as they approach the end of their administration, the corresponding cost of abandonment should increase and they cannot shift the blame to the previous government, especially when there has been an “irregular transfer of power”. The variable is configured as a dummy of occurrence of election, replaced in later models by twin dummies of victory of the ruling coalition/change in office.

Degree of openness. This is approximated by (exports + imports) /GDP. A large ratio means a greater impact of a given devaluation on the aggregate price level and thus a greater cost for the policymaker, so it should reduce the motive to abandon a peg. On the other hand, it can increase the cost of a given appreciation of RER and hence necessitate a relief via devaluation, so that the direction of the final causality is ambiguous.

The unemployment rate. High unemployment has been depicted by SGM as a strong motive to follow an expansionist policy of Keynesian type to stimulate demand, so it should be associated with occurrence of crises. On the other hand, if authorities apply neo-classical policies, unemployment may be addressed with measures that reduce labour cost, such as the abolishment of minimum wage or the liberalisation of the labour market. Extending the argument, an auxiliary army of unemployed can be considered the vehicle to limit the claims of labour syndicates. It is not rare that markets react positively to massive lay-offs of troubled firms, viewing that as a containment of competitiveness erosion. So, it should not be a surprise if, in conjunction with political developments and the presence of high wages and/or inflation, the unemployment variable comes with a negative sign.

Wages. This variable lies at the core of the Obstfeld (1994) analysis; it reflects the inflationary expectations of economic agents that interact in the game-theoretic determination of equilibrium prices and thus the exchange rate. Even if actual and expected growth of the economy is high, an even higher rate of wages' growth may be deemed by the market as an erosion of competitiveness that will, sooner or later, be addressed with a devaluation. In a later model the variable is replaced by the change of the Consumer Price Index (CPI), to test whether direct measurement of total inflation is more relevant. Then, the CPI is maintained along with the unemployment rate in order to embrace potential trade-offs between the two of the type of a Keynesian Phillips curve. However, if estimation is conducted on a period of stagflation, i.e. high inflation along with rising unemployment and plunging production, such as the aftermath of the two oil crises, a fall in wages does not necessarily improve competitiveness. On the contrary, it can lead to a weakening of demand that drags the economy into a downward spiral of negative growth. Then the impact of the variable can be negative.

Contagion. A dummy that takes the value of 1 if a crisis occurs at the same month in any other country within the sample and 0 otherwise. Occurrence of "crisis" is defined according to the particular dependent variable used in each model. Making the measure regional, as in other studies, by assigning countries to several geographic areas and signalling "1" accordingly, would have little meaning in our panel. All countries in our study are classified as developed and markets and institutions treat them as, more or less, similar. The variable should capture pure herding as well as all three aspects of "structural" contagion explained in SGM, namely: (i) trade links (competitive disadvantage from a devaluation of an important partner or competitor, as in Gerlach and Smets, 1995); (ii) macroeconomic similarities (markets speculating that countries with similar economic structure and problems will react in like manner to an attack, as in Buiter et al, 1996) and (iii) financial links (variations in investors' appetite for currency risk without any changes taking place on the fundamentals due to financiers having positions in other, devalued currencies).

We propose that the logic of accumulation applied to the RER can be extended to the rest of the variables. It is plausible to test whether at least some of the regressors have a significant impact only when their misalignment is protracted over a longer period and/or different variables work in different horizons prior to an attack. Some theoretical models offer examples of influences, which work only when accumulated, such as a gradual built-up of financial excess or external sector deficits. Obviously this technique costs a lot of lost observations; thus its use can only be indicative.

2.3 *Composition of the sample*

The database assembled is the largest and most comprehensive in the crisis literature to date. It comprises 11,316 monthly observations of several variables: 23 countries in time series of 40 years, extending from 1960 to 2000. The sample includes all the nations classified in IMF's publication International Financial Statistics (IFS) under the subgroup Industrial Countries.[4]

This study utilises monthly data. Prior approaches had used quarterly or even annual data, not only for reasons of availability of data, but also because the longer-term causalities of crises with fundamentals are clearer in lower frequency. Frankel and Rose (1995), *inter alia*, showed that relationships between the exchange rate and fundamentals weaken substantially in higher frequencies due to noise. However, low frequency eliminates the usefulness of the approach as a predictive tool. Furthermore, examining if presumed relationships, showed to hold in the longer run, are obscured in the short-run due to the operation of the market is a purpose in itself. If even our heteroskedastic and random effects models fail to overturn the phenomenon of frequency-dependent fit, this is evidence that speculative bubbles disrupt exchange rate arbitrage in the short-run. Lastly, the ability of monthly data to capture short-duration attacks, especially the unsuccessful ones, allows a comprehensive study of all occasions of speculative pressure instead of just extraordinary crashes.

The impact of variables was explored in several time horizons, ranging from contemporaneous to one month lagged, jointly with current values or by themselves, and also on a cumulative basis. We make limited use of moving averages in order to avoid generating serial correlation. The "predictive" model using only one-period lagged variables can help to differentiate genuine leads from effects of the attack itself. It can also address the possibility of non-synchronous acquisition or processing of the

[4] The data set draws from sources like the IMF, OECD, Eurostat, Keesing's Record of World Events etc. Details can be found in: http://www.lboro.ac.uk/departments/bs/research/responce-models.html

relevant information from market agents. For a few indicators that are unavailable in monthly periodicity in a few particular years, we employ their quarterly counterparts and repeat the values for 3 consecutive months. Flows are apportioned across the months. The underlying assumption is that agents use the last piece of publicly available information in order to form expectations and decide their action. The reduced variability resulting from the repetition is far outweighed by the wealth of monthly updated information in a host of variables.

We adopt the "exclusion window" technique applied by all existing studies. Observations immediately preceding and following the "crisis" observations are excluded in order to prevent double counting of each episode. To avoid the sample becoming highly unbalanced, thereby drowning any causal relationship, the same is applied to the 'tranquil' (non-crisis) periods as well, which are then used as the control group. However, the use of monthly data leads to a significant imbalance remaining; caution is needed not to attribute this to short-run speculative bubbles. In addition, exclusion windows deliberately create biases in favour of "crisis" observations, hence the models' likelihood estimates should not be interpreted as exact probabilities of an attack.

Most variables enter the estimation in the form of differences of natural logarithms. Details on their configurations are described subsequently in the exposition of each model.

3 The Binomial Benchmarks

As a first stage, we estimated binomial LDV models, which implicitly assume homogeneity of causal relationships between crises and fundamentals. A full report is provided in Anastasatos and Davidson (2004). Overall, the key findings, although revealing some important regularities, increase the level of scepticism on the homogeneity hypothesis as they sketch a picture of noise and low performance, which is increased by divisions of the sample. They provide the rationale for conducting a formal investigation of structural differences among crises and, consequently, adopting multinomial LDV models for optimising estimation and prediction.

A graphical analysis juxtaposing the behaviour of variables in periods of crisis and tranquillity revealed some systematic links of crises with fundamentals. However, in the immediate months surrounding crises, there is little movement of these variables. This fact is bound to make prediction of exact timing of crises very hard. Subsequently, econometric analysis was conducted; results and sensitivity tests are discussed in the exposition of ordered models, as they are analogous.

We also addressed the Flood and Marion (1998) critique that the use of extreme values of the index to signal an episode may lead to loss of many predictable crises. This would happen if interest rates start rising and reserves start dropping before the attack due to uncertainty and a longer-term interest rate applying (Krugman's initial model assumed zero-maturity interest rate). We devised an alternative signalling rule to account for the possibility of crisis jumps being somehow allocated in a number of periods before the attack. More specifically, speculative pressure was approximated as a weighted average of three consecutive observations of the EMP index exceeding a set limit. This specification is enabled by the aptitude of monthly data to capture short and medium-run dynamics of the built-up to a crisis; quarterly or annual data would render this lagging construction useless. However, results proved that, although the fit of the models improves, crises that have a lengthier build up are not necessarily more predictable in terms of their relation with given fundamentals.

Formal specification tests offered a preliminary documentation of the dissimilarity of various episodes. The hypothesis of temporal stability was rejected; in particular, there was a structural break in the point of unification of Germany in 1989. It seems that recent crises are driven primarily from international factors while earlier episodes are more closely associated with traditional domestic fundamentals. We also tested whether safe-havens, i.e. currencies consistently under-performing their multilateral forward rate for a number of years, are likely to experience capital inflows when others are attacked. The hypothesis of cross-sectional stability could not be rejected, although "safe havens" seem to differentiate from other countries and to have also stronger links with domestic imbalances. Comparison with studies in lower frequency and smaller panels showed an increase in inter-temporal heterogeneity. Random effects did not improve classification accuracy. We also detected significant time-dependent heteroskedasticity in our panel and found that taking account of it improved performance considerably.

4 Ordered Models

In the light of the results from the binomial models, we now turn to the formal testing of the similarity of crises of different scale. The benchmark is the ordered *Model 1*, presented in the first column of Table 1. Variables enter in differences of natural logarithms scaled against German values. The model uses the country-specific weighted EMP index of (6) with an 1.5 and 2 standard deviations above the mean thresholds for defining "episodes of lower size" and "major episodes" respectively.

Before proceeding with estimation and analysis, the non-violation of the Parallel Regressions Assumption (PRA) must be established. For an informal assessment, we compared *Model 1* with two equivalent binomial models utilising identical specifications for regressors and y* but employing the various τ_j's of the ordinal model (1.5 and 2 sd above the mean) as their cut-off points. We found that sizes of coefficients of the ORM are very similar to those of the binomial model employing the lower threshold but noticeably different from those of the binomial model with the higher threshold. The divergence is more apparent for the real growth, capital controls, inflation, RER, government victory and contagion variables. Interestingly, longer-term macroeconomic fundamentals, such as growth, were more important for higher scale episodes. Smaller scale episodes on the other hand, seem to be the vehicle for the relief of the economy from an inflation-driven competitive disadvantage.

Subsequently, we formally tested the PRA by Wald tests for the equality of coefficients among models. The Wald statistics are constructed by imposing the equality restrictions onto the coefficients of the ordinal model. The hypothesis of equality with the coefficients of the ordinal model could not be rejected for neither of the binomial models, so that PRA is satisfied. Similarly, PRA was satisfied for the rest of the ORMs; thus, estimation is valid. Nevertheless, some statistics approach the critical value of rejection so that respective tests cannot be considered decisive and a note of caution applies to results. This feature also provides the rationale for the estimation of the Multinomial models, which do not impose the ordinality structure on the data and thus are not limited by the PRA.

Model 2 is estimated on unscaled data, with variables entering in the form of differences of natural logarithms, but it maintains *Model 1*'s setting of threshold values. *Model 3* employs cumulative variables. By the nature of variables, thresholds have to be lowered now, so "lower-size episodes" and "major episodes" are signalled if an observation exceeds the mean by at least 1 and 1.5 standard deviations respectively. ORMs capturing exclusively devaluations were also constructed. *Model 4* employs the dual rule of (7) for capturing currency plunges but it divides them in "minor" and "major" devaluations and vests the criterion with an ordinal structure. Analytically, "1" and "2" observations are defined as:

$$\begin{cases} \text{Index} = 2 & \text{if } y_{it} > \bar{y}_{it} + 1.5\sigma_i \quad \text{and} \quad (y_{it} - y_{i(t-1)})/y_{i(t-1)} > 10\% \\ \text{Index} = 1 & \text{if } y_{it} > \bar{y}_{it} + 1.2\sigma_i \quad \text{and} \quad (y_{it} - y_{i(t-1)})/y_{i(t-1)} > 5\% \\ \text{Index} = 0 & \text{otherwise} \end{cases} \tag{9}$$

Model 4b follows the same logic but one more category is added to signify "crashes", so that devaluations are allocated in three categories:

$$\begin{cases} \text{Index} = 3 & \text{if } y_{it} > \bar{y}_{it} + 1.75\sigma_i \quad \text{and} \quad (y_{it} - y_{i(t-1)})/y_{i(t-1)} > 15\% \\ \text{Index} = 2 & \text{if } y_{it} > \bar{y}_{it} + 1.5\sigma_i \quad \text{and} \quad (y_{it} - y_{i(t-1)})/y_{i(t-1)} > 10\% \\ \text{Index} = 1 & \text{if } y_{it} > \bar{y}_{it} + 1.2\sigma_i \quad \text{and} \quad (y_{it} - y_{i(t-1)})/y_{i(t-1)} > 5\% \\ \text{Index} = 0 & \text{otherwise} \end{cases} \tag{10}$$

In both *Models 4* and *4b* regressors enter in differences of logarithms scaled against German values and the deutschemark cross-exchange rate, serving as the latent y*, is country-specific standardised.

Comparative analysis of all ordered models verifies that the main causal relationships between speculative pressure and fundamentals uncovered in their binomial counterparts reiterate in a robust manner. Only the significance of less relevant variables alternates in the various models. The main driving forces are the RER, excessive money supply and inflation, contagion, weak real growth and the current account. Contagion is highly significant; however, the evidence is not conclusive as the variable may reflect unobservable shocks, common to all countries. Some weaker evidence is also provided in non-reported models for unemployment. The significance of 'Deficit' seems to have been limited by the availability of international credit lines. 'Openness' seems to reduce the motive to abandon a peg by magnifying the cost of a given devaluation on the aggregate price level but the pattern is not stable as it reverses in other models. Capital controls seem to be an important deterrent if imposed early but positively correlated with crises otherwise. It is also interesting that the private loans variable was insignificant in all estimated models. To the extent that the proxy captures financial weakness, it shows that the informational asymmetries framework is applicable to shallow and illiquid markets and not to the developed countries of our sample. This reinforces the heterogeneity hypothesis.

An important result contradicting previous studies in lower frequency is that, in all estimated models, binomial and multinomial, the RER appears with a negative sign, unless cumulative. It seems that appreciation, in the short-term, leads the market to expect even more appreciation, rather than mean reversion of the RER. Conversely, when devaluation starts, it lasts for a few months on average, as evident from visual inspection of data, and hence the negative sign. So, short-term dynamics and the psychology of the market reflect a steady-as-going expectation. In the longer term, if accumulated to a considerable degree and duration, real appreciation is a significant driving force behind crises. This result validates the lagged mean-reversion hypothesis that had gained credibility from findings in previous studies.

Sensitivity analysis revealed that, although some basic findings are sufficiently robust, results are influenced to some extent by specification. Overall, sensitivity cannot be attributed to the ordinal specification as there is more of it among different specifications of y^* and **X**, than among binomial models and their ordinal counterparts. Three sources of sensitivity can be traced: (i) the definition of a crisis, as reflected in the construction of the limited-dependent variable, (ii) the composition of the sample used in estimation, particularly where model requirements have discarded certain observations, and (iii) the high frequency of data. The remaining sensitivity though should be attributed to inherent dissimilarity of examined episodes.

More specifically, changing the set of explanatory variables as described in section 2.2 has limited impact on results. Exclusion of consistently insignificant variables from models results in the repetition of the same patterns enhanced rather than weakened; tests confirmed their irrelevance. Accumulated variables depict more clearly causalities described by theoretical models but fail to improve classification. Lagging variables resulted in dramatic fall of the fit and classification scores. This is in contrast with the findings of Frankel and Rose (1996) that, with annual data, lagging strengthens relationships, and it shows the vital impact of data frequency on results. Models employing the criterion of (7), i.e. capturing only actual devaluations, perform better than models employing EMP indices. This hints that repelled attacks are harder to capture and explain. Among EMP indices, the country- weighted index (6) performs better but none improves noticeably in classification; scaling against German values does not change results significantly.

The AIC suggested that the fit of the ORMs is slightly deteriorated in comparison to binomial models that follow the same specification of explanatory variables. Indeed, the more ordinal structure is imposed onto the data, the higher the AIC becomes, indicating a worse fit. This is a natural consequence of the development of the models, bearing in mind that the transformation of the continuous y* into limited-dependent variables is an *ad hoc* construction. Then, a finer categorisation of episodes further obscures the link between the end dummy and its continuous benchmark, and leaves mores space for misclassification. Furthermore, if the satisfaction of the PRA is ambiguous or unattainable, this should be reflected in performance measurement.

Therefore, the most interesting finding is in respect of the accuracy of classification, as it is closely related to the heterogeneity question. The models are quite successful in correctly classifying larger crashes, even if some are misclassified in the lower category, but its performance in capturing smaller scale episodes is poor. This should be no surprise. Intuitively, episodes of larger scale should include more pronounced misalignments of fundamentals if any causal relationship exists. In smaller episodes, especially repelled attacks, it is empirically difficult to distinguish between crises and tranquillity. This finding complements the tendency of binomial models to perform worse when the threshold is set lower. Then, a broader range of episodes of smaller and larger scale is captured; this reduces the fit of models and attaches some sensitivity to the coefficients. It follows that setting the threshold of the ORM's lower category sufficiently high would increase the percentage of correct calls. However, this practice limits the usefulness of the models as a guide for detecting the occurrence of less than catastrophic crises.

4.1 Heteroskedastic and Random Effects Ordered Models

In order to conduct the LM test for heteroskedasticity, all regressors of the respective models enter the skedastic function. Results of the LM test for *Models 2-4b* are showed in the foot of Table 1. As seen in section 3, the homoskedasticity hypothesis for all models is rejected convincingly, even at the 97.5% level. To deal with this, we estimate heteroskedastic ordered probits using Harvey's (1976) specification. Variables with sizeable standardised variance terms that approach significance are included in the skedastic function, up to the extent that allows convergence as indicated in Table 2. The heteroskedastic ORMs exhibited in Table 2 relate to the numbering of the equivalent simple ordinal model in Table 1 but with the letter *h* added. These heteroskedastic models have behaviour comparable to their simple ordered counterparts but with improved fit as indicated by the AIC measure, the reduction in the size of the coefficients of significant variables attesting to the importance of allowing for heteroskedasticity in the model specification. Overall, the correct in-sample calls improve noticeably over those generated by the simple univariate ordinal models, the heteroskedastic specification improving the classification performance of all ordered models by 10-25% by exploiting the time-series features more effectively. These heteroskedastic models correctly called around 25-50% of all crisis instances - a fairly good outcome given the increase in noise associated with using higher frequency data- but still poor in prediction terms.

The random effects models showed that month-specific idiosyncrasies exist. However, in contrast to heteroskedastic models, they did not improve classification accuracy. Whether it is heteroskedasticity *per se* that was addressed or some other specification flaw, e.g. neglected non-linearities, it is not easy to say. Again, even after the improvement, prediction remains mediocre. Thus, episode dissimilarity remains the prime suspect cause of the remaining faults.

5 Multinomial Models

We now proceed to the estimation of Multinomial Logit models. In *Model 5* of Table 3, a "successful" attack is postulated to have occurred when (i) the country-specific weighted EMP index (6) with a threshold of 1.5 standard deviations above the sample mean signals an instance of crisis, of whichever sort, and (ii) at the same month, an abnormal devaluation of considerable magnitude has occurred in the country. The latter requires the fulfilment of a dual criterion:

$$\begin{cases} \text{Index=1} & \text{if } y_{it} > \overline{yit} + 1.5\sigma_i \quad \textit{and} \quad (y_{ti} - y_{i(t-1)}) / y_{i(t-1)} > 4\% \\ \text{Index=0} & \text{otherwise} \end{cases} \tag{11}$$

If the EMP index signals "1" but the devaluation dummy does not, this is said to be a "failed" attack. The particular indicators have been chosen because they both employ the country-specific weighted deutschemark exchange rate in the composition of their latent y*, so they are directly comparable. Also, the criterion of the devaluation dummy is strict enough to allow a lucid segregation of crises. Variables enter in differences of natural logs scaled against German values.

A variation of the model, *Model 5b*, employing a stricter criterion for identifying crises, is also estimated. In *Model 5b*, a "successful" attack is postulated to have occurred when the country-specific weighted EMP index (6) is at least 2 standard deviations larger than the sample mean, *and* the dual criterion signals an abnormal devaluation of considerable magnitude according to:

$$\begin{cases} \text{Index=1} \quad \text{if } y_{it} > \overline{yit} + 1.75\sigma_i \quad \textit{and} \quad (y_{ti} - y_{i(t-1)})/y_{i(t-1)} > 15\% \\ \text{Index=0} \quad \text{otherwise} \end{cases} \tag{12}$$

If only the first condition is satisfied, the attack is classified as "failed". The specification of informational variables is identical to that of *Model 5*. The two models are comparatively presented in Table 3, apparently with a different coefficient vector for each outcome.

The two outcomes present remarkable dissimilarities among them. Importantly, since coefficients for different categories come from a single model, estimated on the same data set, the possibility of sample selection driving any differentiation due to missing values can be ruled out.

The fit of the models, as measured by the AIC, is stronger than that of equivalent binomial models, with the best performer again being *Model 5b* with the stricter rule for detecting instances for study. However, classification accuracy does not improve as a whole just as a result of the stricter categorisation of crises. The imposition of more structure onto the data, even though not ordered but of quasi-qualitative nature, obstructs correct classification and predisposes in favour of "tranquillity". Overall, episodes ending in a major devaluation are more readily correctly classified than 'repelled' or 'minor' attacks. For an illustration, the classification table of *Model 5* is reported on the foot of Table 3.

Next, we construct the LR test for multidimensionality[5]. The result is LR =20.3067; with a 95% critical value of the χ^2 distribution for 10 degrees of freedom of 18.31, we can reject the null of indistinguishability. Therefore, the use of the MNLM is justified. However, this result is much more than technical justification for the use of the multinomial platform of models for unordered outcomes. It is formal evidence that not all crises are alike. If episodes classified in various categories, on whichever criterion, cannot be described by a single equation then there are systemic differences amongst them. Therefore, the application of a uniform model without proper consideration of the length and span of the panel data on which it is estimated is destined to fail. The result is also in

[5] The government victory dummy had to be excluded because its limited variability was blocking estimation. This should not lead to a false rejection of indistinguishability. The variable has limited significance anyway.

contrast with the finding of Eichengreen et al. (1995), who report only few significant differences when they separate successful and unsuccessful crises by using *ex post* and *ad hoc* criteria.

In analysing the economic implications of the results, economic and political fundamentals are, as expected, more strongly associated with successful attacks. However, the contagion variable loses significance in the instances of attacks that force substantial devaluation. It has to be assumed that in order for an episode of more than temporary importance to occur, fundamental domestic imbalances have to exist separately from international transmission.

An important result is that poor expectations of real growth, as well as the objective lack of it, are associated with successful attacks. This fact verifies that expectations are important but also that they tend to focus on longer-term considerations too. It is reasonable then to assume, that any growth problems existing before the attack will be aggravated by both the prevailing pessimism in the market and the objective effects of the crisis per se. Rising interest rates and turbulence in markets, possibly accompanied by inflationary pressures if the attack has succeeded in under-valuing the exchange rate, are certainly adverse factors for the recovery of investment activity. The afflicted country will pay a high cost in terms of output lost and possibly a setback of trust in transitory and adjustment policies. Also, it is confirmed that absence of capital controls favours success of attacks. This result has not been clearly established in previous studies. The occasional positive sign of the variable has to be attributed to recurring attacks in which limitations in the capital account have been imposed *post factum*. Inflation, if not the sign of more structural imbalances, and if not in excess of the rate of devaluation allowed by the exchange rate regime in place, is associated with more ephemeral turmoil.

The qualitative separation of currency episodes allows the unemployment variable to approach significance with a negative sign, a finding that is absent in other models. This is most interesting because it comes from the exact sample that did not render a similar result in binomial models, therefore it cannot be considered a peculiarity caused by sample selection. The stronger negative impact of unemployment on failed attacks is a sign that markets do perceive unemployment as a motivation for competitiveness-boosting policies but this perception is not crucially linked with macroeconomic causalities that lead to a devaluation. If relevant at all, unemployment does not lead to the reduction of labour cost because of market rigidities, or the labour cost itself is not decisively related to competitiveness. The argument is propped up by the strong negative impact of the lack of real growth in the same model.

Model 6 provides a cross-checking of the ORM results on separation of episodes according to scale, while addressing the concerns on the validity of the ORM, due to the ambiguous satisfaction of the PRA. MNLM is a legitimate alternative for cases in which the strict ordinality of outcomes is in doubt. Therefore we use it to re-estimate *Model 1*. Identical specification of both regressors and regressand is followed in order to establish equivalence. It can be seen that in general terms the model repeats the results of *Model 1*, although the size of the coefficients cannot be directly compared between probit and logit models. AIC records a fit very similar to that of *Model 1*. In terms of correct in-sample calls, *Model 6* correctly classifies a couple of incidents in excess of its ordinal counterpart, still without escaping mediocrity. However, the qualitative nature of the distinction among the various "crisis" outcomes in the MNLM allows the elevation of some causal relationships. Macroeconomic measures that are considered to be more fundamental and with longer-term influence tend to be more strongly associated with larger scale episodes. The prominent example is real growth, which has a dramatic difference in the two categories and a strong negative impact only in "larger scale" episodes. To a lesser extent the same holds for current account too. Inflation remains only relevant to minor incidents, possibly recurring attacks; perhaps its effect associated with structural imbalances is picked up by other variables. The same holds for the dummy for government victory; it is obvious that not all political developments affecting crises can be captured by this variable.

We also conduct the LR test for indistinguishability of the two "crisis" outcomes with respect to the variables on *Model 6*. Since some variables have been omitted from **X** in this model, we would expect the test to have less power. Nevertheless, the LR statistic is calculated to be LR=15.75994. The 95% critical value from the χ^2 distribution for 9 degrees of freedom of 16.92; the 90% critical value is 14.68. Therefore, multidimensionality cannot be rejected with certainty for this model too.

6 Conclusion

This study offers strong evidence of structural dissimilarity between speculative attacks, implying that the inherent hypothesis spanning most empirical studies that all crises are driven by the same imbalances and follow the same process is misguided. Some patterns do emerge from empirical results across specifications, irrespectively of whether involved fundamentals work deterministically or indirectly as a co-ordination device for expectations. Our results are consistent with previous findings in that M1/P and its mirror image, inflation, together with the RER are the most consistent determinants of crises. Importantly and in contrast to previous studies, appreciation appears with a negative sign, unless cumulative. This result is consistent with the lagged mean-reversion hypothesis. Weaker evidence is also provided for the current account, unemployment, budget deficit and the lack of real growth. The existence of capital controls is a significant deterrent of crises but not if imposed *post-factum*. Contagion is important and it complements or even substitutes domestic imbalances, especially so in more recent crises, which occur in an increasingly inter-related global environment. On the other hand, wages, credit growth and loans to the private sector are consistently irrelevant. The results are fairly robust across different configurations of the data and explanatory variables, but fit and correct calls vary in relation to how a "crisis" is defined and approximated. This fact was played down in most previous studies. But this is not responsible for our finding on heterogeneity.

What are the sources of heterogeneity? Some of it may have come from the inclusion of voluntary devaluations in gleaned episodes, although our use of EMP indices should have limited this danger. Also, mediocre predictive performance of the models verifies that the causal relationships of crises with fundamentals are obscured in higher frequencies from greater trading noise. This makes forecasting difficult, despite using more up-to-date data. Even if early warning signals can be traced, those could work as an update for the co-ordination mechanism of expectations in a game-theoretic context, thus endogenising predictability. Thus, the assertion of Frankel and Rose (1995), that the role of fundamentals is present in the determination of the exchange rate but limited by speculative bubbles in the short-run, gains support. This implies that systemic differences amid attacks on various exchange rate regimes may arise from differences in regimes' volatility. This is more likely since the majority of observations in our sample come from free or managed floats (many pre-Bretton Woods observations were missing). Based on the findings of Eichengreen et al. (1995), we could possibly improve the fit of models by excluding the less predictable component of episodes, i.e. regime transitions. Nevertheless this is not a sound theoretical choice since there is no way to foresee *ex ante* whether an attacked peg will be realigned, abandoned, or successfully defended.

If bubbles exist, multiple equilibria cannot be ruled out. However, there is no apparent way to ensure that all relevant macroeconomic and political measures have been correctly modelled, which is a prerequisite before admitting herding behaviour. The persistent significance of the contagion variable reinforces the last dilemma. We can only repeat the Blanco and Garber (1986) postulation that 1st and 2nd generation models are observationally equivalent when expectations for fundamentals change for good reasons, not observed by the econometrician. Whether crises that escape prediction are caused by an alternative set of fundamentals or self-fulfilling prophecies, both eventualities constitute heterogeneity among crises. Despite those limitations, we have been able to establish by formal tests that such heterogeneity exists. We found that crises relationships with fundamentals do differentiate according to the crisis' scale, upshot and era of eruption. This is inherent dissimilarity of episodes and proves that crises cannot be treated in a unified manner.

Crises of a larger scale, as well as the most successful attacks, tend to be associated with more structural factors like the lack of real growth and unemployment. More ephemeral episodes on the other hand had an overall looser link with fundamentals. Our use of in-sample classification criteria allows the exploitation of this result for prediction of future incidents. There is no reason to believe that crises do not differ according to other criteria too, not easily quantifiable, such as the exact exchange rate regime upon which they occur or institutional factors of the economy.

The use of monthly data in this study accurately captured the dynamics of a gradual decline: devaluations are the culmination of a lengthier process of adjustment rather than a digression on the spur of a moment. Even so, cumulative influences cannot predict the exact timing of an attack. It proved difficult to temporally identify a currency episode with a single extraordinary movement of explanatory variables, even when the entire process is taken into account. This can only mean that the eruption of crisis is not associated with crossing an obvious deterministic threshold. This is corroborated by the fact that levels of variables had less explanatory power than their dynamics.

It follows that an indiscriminate application of a general model built on the findings on random past crises cannot achieve a dramatic increase of the percentage of correct calls under *any* specification. In other words, a judicious choice of the temporal and cross-section spread the sample is of equal importance to the methodology itself for successful prediction of future episodes.

The international investor can gain valuable insights from these results but the development of formal, coherent rules aimed at yielding superior hedging performance is inappropriate. Our results are most valuable to the policy maker, although not in a fashion of early warning signals but rather as an instrument of tracing underlying processes. The multi-dimensionality of the crises-fundamentals relationship means that the general exchange rate strategy should combine corrective measures for fundamentals visibly incompatible with the desired level of the exchange rate and actions aimed in soothing market sentiment, so that crises unjustified by fundamentals could be avoided.

The notion of macroeconomic determinism as the rigid framework of crises seems less appropriate than the effort to detect potential weaknesses that can focus the attention of global markets to a particular economy. The increasing mobility of international capital and its inherent propensity to scrutinise speculative opportunities or hedging necessities, tend to expose and magnify weaknesses that would otherwise be internally solvent. There is every reason to believe that this tendency will be enhanced in the future.

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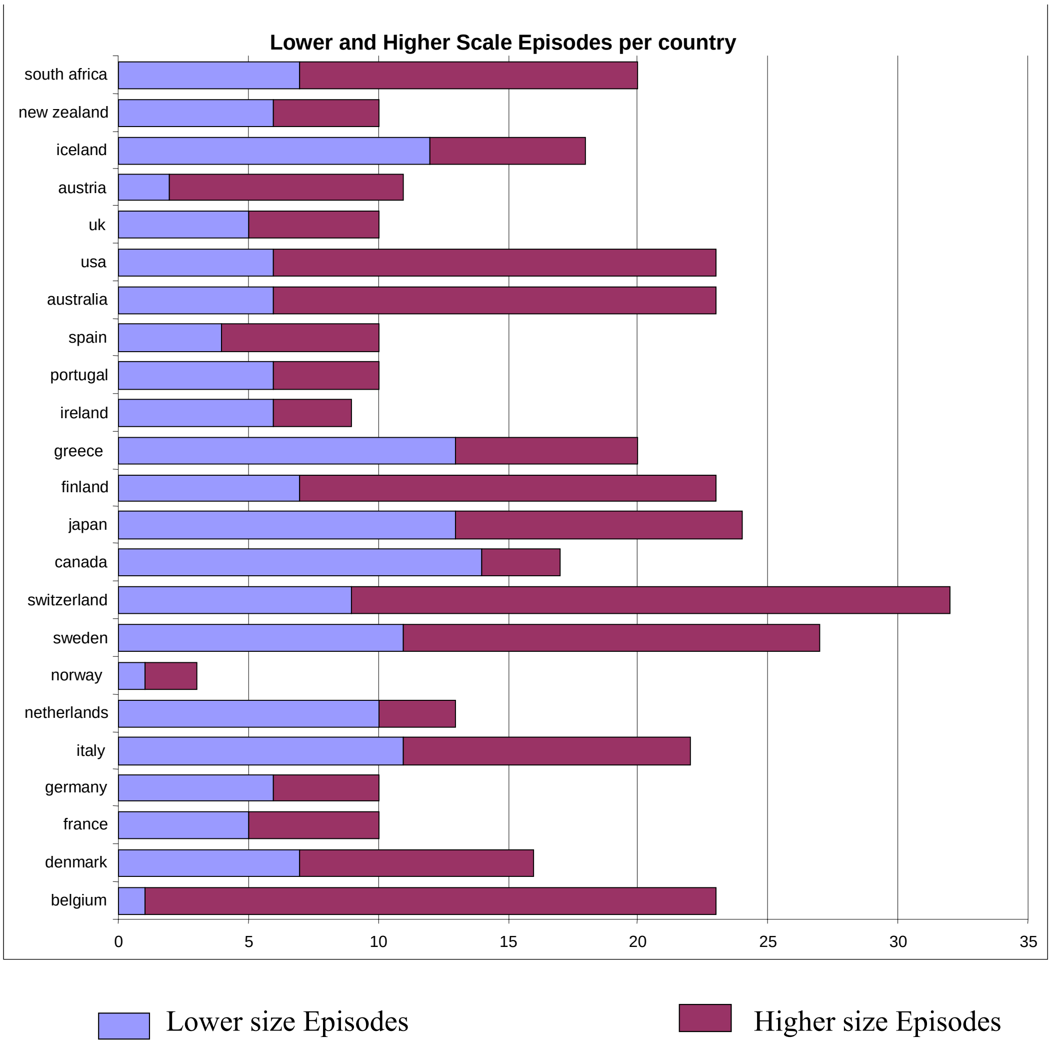
Figure 1: Geographical Allocation, Lower and Higher Scale Episodes



Episodes gleaned from $EMP_{i,t} = [(s_{i,t}/3\sigma^{s}_{i}) + ((i_{i,t} - i_{G,t}) / 3\sigma^{int}_{i}) - ((\% \Delta r_{i,t} - \% \Delta r_{G,t}) / 3\sigma^{r}_{i})]$ as : "smaller crisis" if: $2.0\sigma+\mu >$ obs $>1.5\sigma+\mu$, "major episode" if: obs$>2.0\sigma+\mu$

Table 1: Ordinal Models	Model 1	Model 2	Model 3	Model 4	Model 4b
Constant	**-18.990** (.0000)	**-16.342** (.0000)	-.632 (.5271)	**-14.456** (.0000)	**-14.238** (.0000)
Capital controls	-.568 (.5702)	1.503 (.1329)	-.056 (.9555)	**-2.293** (.0218)	**-2.307** (.0210)
Election		.020 (.9844)		.957 (.3387)	.711 (.4769)
Government Victory	.819 (.4126)		.661 (.5086)		
Contagion	**4.757** (.0000)	**5.478** (.0000)	**3.975** (.0001)	**2.138** (.0325)	**2.096** (.0361)
Current account	**-1.892** (.0585)	-.831 (.4062)	-.591 (.5545)	-1.028 (.3039)	-1.316 (.1881)
M1/P		**4.145** (.0000)	.214 (.8303)	1.260 (.2075)	.905 (.3657)
Deficit	-.045 (.9645)	-.126 (.8999)	.222 (.8240)	-1.352 (.1765)	-1.476 (.1400)
Shares index		-.212 (.8319)	-1.423 (.1548)	-.496 (.6202)	-.394 (.6938)
Unemployment	-.562 (.5741)	.578 (.5630)	**2.065** (.0390)	.170 (.8651)	.208 (.8356)
Wages		.047 (.9624)		-.669 (.5032)	-.781 (.4349)
Credit		-.135 (.8924)			
Private loans		.300 (.7640)			
Openness		-1.317 (.1880)	**-2.002** (.0453)	-.827 (.4080)	-.840 (.4007)
RER	**-9.766** (.0000)	**-11.934** (.0000)	**1.726** (.0844)	**-17.801** (.0000)	**-17.857** (.0000)
Real Growth	-1.444 (.1488)		**-1.989** (.0467)		
Inflation	**2.575** (.0100)		.321 (.7483)		
AIC	0.288173	0.393112	0.472779	0.373614	0.429182
LR test P-value	0.0	0.0	0.0	0.0	0.0
Heteroskedasticity LM statistic		29.23299	52.79402	55.75630	51.94929
χ^2 critical value at 95% (degrees of freedom)		22.36 (13)	21.03 (12)	19.68 (11)	19.68 (11)

Model 1: Ordinal, $EMP_{i,t} = [(s_{i,t}/3\sigma^{s}_{i}) + ((i_{i,t} - i_{G,t}) / 3\sigma^{int}_{i}) - ((\% \Delta r_{i,t} - \%\Delta r_{G,t}) / 3\sigma^{r}_{i})]$,
"smaller crisis" if: $2.0\sigma+\mu >$ obs $> 1.5\sigma+\mu$, "major episode" if: obs$>2.0\sigma+\mu$
regressors: data scaled against German values.
Model 2: Ordinal, estimated on unscaled data, "smaller crisis" if: $2.0\sigma+\mu >$ obs $> 1.5\sigma+\mu$, "major episode" if: obs$>2.0\sigma+\mu$
Model 3: Ordinal, cumulative regressors, "smaller crisis" if: $1.5\sigma+\mu >$ obs $> 1.0\sigma+\mu$, "major episode" if: obs$> 1.5\sigma+\mu$
Model 4: capturing devaluations only, divided in "minor" and "major", data scaled against German values.
Model 4b: capturing devaluations only, divided in "minor", "major" and "crashes", data scaled against German values.
Standardised coefficients reported, significant at 10% showed in bold. P-values in parentheses, i.e. P[|Z| >z].

Table 2: Heteroskedastic Ordinal Models

	Model 2h	Model 3h	Model 4h	Model 4bh
Constant	**-8.764** (.0000)	.083 (.9342)	**-9.431** (.0000)	**-9.555** (.0000)
Capital controls	-1.350★ (.1772)	-.797★ (.4253)	**-2.555** (.0106)	**-2.022** (.0431)
Election	.768★ (.4424)		.479 (.6320)	.887 (.3750)
Government Victory		.253 (.8006)		
Contagion	**2.810★** (.0050)	**3.566** (.0004)	**1.651** (.0993)	**1.777** (.0756)
Current account	-.187 (.8515)	-.376 (.7069)	**-1.887★** (.0592)	-1.171 (.2417)
M1/P	**3.554★** (.0004)	.245 (.8062)	1.290 (.1972)	1.130 (.2583)
Budget Deficit	**-1.865★** (.0621)	.545 (.5859)	-.555★ (.5792)	-.105 (.9163)
Shares Index	-.697★ (.4856)	-1.402 (.1609)	-1.496★ (.1346)	-1.467★ (.1424)
Openness	-1.371★ (.1702)	.559★ (.5760)	-1.115 (.2647)	.052∴ (.9584)
Unemployment	.890 (.3736)	**1.760** (.0785)	.143 (.8864)	.195 (.8457)
Wages	1.321★ (.1866)		-.640 (.5222)	.586★ (.5576)
Credit	.651★ (.5151)			
Private loans	-.364 (.7159)			
RER	**-4.306★** (.0000)	**3.737★** (.0002)	**-8.342★** (.0000)	**-8.850★** (.0000)
Real Growth		-1.592 (.1114)		
Inflation		-.387 (.6991)		
AIC	0.389333	0.468432	0.348506	0.405754
LR test P-value	0.0	0.0	0.0	0.0

Model 2h: Ordinal, heteroskedastic counterpart of Model 2, estimated on unscaled data, "smaller crisis" if: $2.0\sigma+\mu >$ obs $>1.5\sigma+\mu$, "major episode" if: obs$>2.0\sigma+\mu$
Model 3h: Ordinal, cumulative regressors, heteroskedastic counterpart of Model 3, "smaller crisis" if: $1.5\sigma+\mu >$ obs $>1.0\sigma+\mu$, "major episode" if: obs$> 1.5\sigma+\mu$
Model 4h: capturing devaluations only, divided in "minor" and "major", heteroskedastic counterpart of Model 4, data scaled against German values.
Model 4bh: capturing devaluations only, divided in "minor", "major" and "crashes", heteroskedastic counterpart of Model 4b, data scaled against German values.
Standardised coefficients reported. significant at 10% showed in bold, P-values in parentheses, i.e. P[|Z| >z].
Variables entering the scedastic function are depicted with (★).

Table 3: MNLM: Successful vs Repelled Attacks, Larger Scale vs Smaller Scale Episodes

	Model Successful Attacks	5 Failed Attacks	Model Successful Attacks	5b Failed Attacks	Model “smaller” episodes	6 “larger ” episodes
Constant	**-10.104** (.0000)	**-13.683** (.0000)	**-9.006** (.0000)	**-9.514** (.0000)	**-12.314** (.0000)	**-12.370** (.0000)
Capital controls	-1.116 (.2645)	.291 (.7708)	**-2.215** (.0268)	.494 (.6214)	.061 (.9515)	-.788 (.4308)
Government Victory	**2.982** (.0029)	.000 (1.0000)	1.214 (.2248)	.000 (1.000)	**1.648** (.0994)	.634 (.5264)
Contagion	**2.959** (.0031)	**3.571** (.0004)	.783 (.4338)	**3.206** (.0013)	**4.132** (.0000)	**3.771** (.0002)
Current account	-.008 (.9932)	**-2.456** (.0141)	-.570 (.5689)	-1.546 (.1221)	-1.185 (.2361)	**-1.650** (.0997)
M1/P	**2.320** (.0203)	**2.376** (.0175)	1.174 (.2405)	-.431 (.6666)		
Budget Deficit	.693 (.4656)	-.381 (.7033)	.316 (.7520)	-.572 (.5674)	.328 (.7428)	-.362 (.7170)
Growth Expectations	**-1.841** (.0657)	.204 (.8385)	**-2.214** (.0268)	-.778 (.4368)		
Openness	.229 (.8187)	-.544 (.5865)	-.445 (.6563)	.173 (.8628)		
Unemployment	-.668 (.5042)	-.577 (.5638)	-.356 (.7215)	-1.544 (.1225)	-.125 (.9009)	-.833 . (4051)
RER	**-7.723** (.0000)	**-6.540** (.0000)	**-6.457** (.0000)	**-4.036** (.0001)	**-6.267** (.0000)	**-8.069** (.0000)
Real Growth	-1.043 (.2970)	-.364 (.7160)	-1.298 (.1944)	**-2.599** (.0094)	1.145 (.2722)	**-2.196** (.0281)
Inflation	.657 (.5113)	**3.188** (.0014)	-.896 (.3700)	1.112 (.2662)	**2.718** (.0066)	1.003 (.3160)
AIC	0.290964		0.179915		0.292789	
LR test P-value	0.0		0.0		0.0	

Classification table for *Model 5*

Predicted Outcome

Actual Outcome	**Tranquillity**	**Devaluation**	**Repelled attack**	**Total**
Tranquillity	1516	3	3	1522
Devaluation	35	3	2	40
Repelled attack	17	0	8	25
Total	1568	6	13	1587

Model 5: Multinomial, estimated on data scaled against German values. Successful attack if EMP=1 *and* Devaluation dummy =1;Failed attack if EMP=1 *and* Devaluation dummy =0.
Model 5b: Multinomial, estimated on data scaled against German values, stricter criterion for identifying crises.
Model 6: Multinomial, $EMP_{i,t} = [(s_{i,t}/3\sigma^{s}_{i}) + ((i_{i,t} - i_{G,t}) / 3\sigma^{int}_{i}) - ((\% \Delta r_{i,t} - \%\Delta r_{G,t}) / 3\sigma^{r}_{i})]$,
“smaller crisis” if: $2.0\sigma+\mu >$ obs $>1.5\sigma+\mu$, “major episode” if: obs$>2.0\sigma+\mu$,
regressorss: differenced, scaled against German values.
Standardised coefficients reported, significant at 10% showed in bold. P-values in parentheses , i.e. P[|Z| >z].