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Does Central Bank Intervention Influence the Probability of a Speculative Attack? Evidence from the EMS

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Editorial

In this paper, Helmut Stix studies the effects of central bank intervention during the 1992-1993 EMS crises on the D-mark/peseta and D-mark/French franc exchange rate. In particular, it is analyzed how interventions affected the probability of a speculative attack and market participants' expectations about realignments. The findings provide evidence that interventions seem to have increased both the expected realignment rate and the probability of a speculative attack. Furthermore, there is some evidence that this effect arises for publicly known but not for secret interventions.

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Helmut Stix*

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Abstract

This paper studies the effects of central bank intervention during the 1992-1993 EMS crises on the D-mark/peseta and D-mark/French franc exchange rate. In particular, it is analyzed how interventions affected the probability of a speculative attack and market participants' expectations about realignments. The findings provide evidence that interventions seem to have increased both the expected realignment rate and the probability of a speculative attack. Furthermore, there is some evidence that this effect arises for publicly known but not for secret interventions.

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1 Introduction and Motivation

In September 1992, shortly after the lira and the pound left the Exchange Rate Mechanism and close to the French Maastricht referendum, the French franc came under pressure. Intervention by the Banque de France was not able to calm the situation. In reaction, the French authorities announced more intense interventions, which, according to then French finance minister Michel Sapin, should "make the speculators pay ... hit them where it hurts—in their wallets". And, a little bitter, Monsieur Sapin added that "during the revolution, such people were known as agioteurs (speculators) and they were beheaded".[1] Despite this threat, the attacks on the French franc and several other currencies continued. How this episode ended is well known: At the end of July 1993, the authorities had to give in and the bilateral fluctuation bands were widened.[2]

What had happened? Many explanations have been given in the literature that explain why the European Monetary System got under such a strain: macroeconomic imbalances, self fulfilling speculative attacks, political uncertainties, to mention just a few. However, the role of one very important policy instrument—interventions—has gained little attention so far. This paper presents new results about the role of intervention in the EMS during this period of speculative pressure. In particular, the paper discusses the empirical implications of two speculative attack models. It is argued that a Markov switching model provides a suitable framework for analyzing the impact of intervention not only for empirical reasons, which is in line with several papers that analyze periods of speculative attacks in the EMS (e.g. Martinez-Peria, 2002), but also from a theoretical perspective. Then, by employing data on daily intervention amounts by the Banco de España, the Banque de France and—in a limited sense—the Bundesbank in the period from June 1992 until the end of July 1993, the effectiveness of intervention is analyzed. In particular, it is tested whether intervention was able to decrease the probability of a speculative attack. Additionally, we estimate

[1] All statements made by Michel Sapin. Source: Reuters News, 'FRANCE: France, Germany win round in battle to save EMS', September 23, 1992.

[2] In 1992 the fluctuation bands was $\pm 2.25\%$ except for Spain and Portugal with bands of $\pm 6\%$. On August 1, 1993 the bilateral fluctuation bands were widened to $\pm 15\%$.

the impact of intervention on the expected realignment rate similar to Koedijk, Mizrach, A. & de Vries (1995) who study the pre 1992 period of the Belgian Franc/German Mark exchange rate. Apart from the evaluation of the effectiveness of intervention during this particular time period and institutional arrangement, it is tested whether the effectiveness of intervention is different between publicly recognized and secret interventions.

The last empirical question is motivated by the well known fact that many central banks tend to keep their intervention activities concealed.[3] In the case of the European Monetary System, an interesting hint for explaining the secrecy of intervention can be found in a statement of former Bundesbank president Helmut Schlesinger. By referring to the experience of the September 1992 crisis, he said that "the way in which central banks had been obliged to support weak European currencies was a *powerful incentive for speculation*".[4] In a similar spirit, Sarno & Taylor (2001) have recently conjectured that central banks use secret interventions in order to avoid that intervention triggers a speculative attack. In particular, if market participants are aware of intervention, then the more central banks intervene to defend the peg, the higher will be the transfers to speculators after the authorities abandon the peg and therefore the higher will be the incentive to attack. Thus, publicly known interventions but not secret interventions might actually increase the probability of a speculative attack. Certainly, the presence of such a perverse effect of intervention would have direct policy implications (e.g. for the EMS II). By analyzing the role of secret versus public intervention, this paper attempts to shed light on the relevance of this channel.

The remainder of the paper is organized as follows: Two speculative attack models and their implications for the effectiveness of intervention are briefly discussed in Section 2. Also, in Section 2 a Markov Switching model is presented and some testable hypotheses

[3]This can be observed across exchange rate regimes. For example, interventions after the Plaza agreement for floating rates and interventions in the European Monetary System before and after 1993 were all conducted more or less secretly. Recently, intervention by the European Central Bank were conducted secretly. Sarno & Taylor (2001) have termed this the "secrecy puzzle".

[4]Reuters News Service, "UK: ERM turmoil - from Danish No to Edinburgh summit", December 7, 1992. Italic font added by the author to highlight that the italic phrase itself is a citation.

derived. In Section 3 stylized facts of the exchange rate and intervention data are discussed. The estimation results of the Markov switching model are presented in Section 4. In Section 5 it is analyzed how interventions affected market participants' expectations. In Section 6 all results are discussed and tests for reverse causality are presented. Concluding remarks and some directions for future research are given in Section 7.

2 Does Intervention Influence the Probability of a Speculative Attack?

In this section, a model by Isard (1995) and a model by Obstfeld (1994) will be briefly and informally presented. It is then discussed how the two approaches can be combined to predict the above mentioned perverse effect of intervention. Basically, the idea of combining both models rests on work of Flood & Marion (1999) and Sarno & Taylor (2001). In contrast to these authors, we emphasize the methodological implications of the two models. These implications will be exploited in choosing an empirical model.

2.1 Speculative Attack Models and Intervention

The rules of European monetary integration gave countries the option to leave the integration process during stage II. Models that analyze such escape clause arrangements, the resulting incentives for governments and its economic consequences are therefore particularly appealing in the context of the 1992-1993 European Monetary System crises. One such model has been proposed by Isard (1995). In this model it is assumed that the government can opt to devalue if a large negative output shock hits the economy. By doing so, the government stimulates the economy because of the presence of short-run nominal rigidities. When evaluating its options, the government compares the loss it incurs when staying with the peg with the loss it incurs when leaving the peg. Trading off the rule (the peg) versus discretion (the decision to devalue) is not costless. In particular, the policy maker has to consider two types of costs. First, by abandoning the peg the government looses credibility. This loss can be interpreted as fixed costs to society. Second, a realignment will alter the valuation of foreign exchange reserves. Hence, if the government intervenes (a change in the stock of reserves) and then realigns (a change in the valuation of this stock), it will incur foreign exchange valuation changes.[5]

If, as assumed by Isard (1995), agents have rational expectations and full knowledge of both the loss function and the amount of intervention, then agents will take these (valuation) costs into account when forming realignment expectations.[6] The extent of the government's exposure to valuation losses in case of a realignment, signals the commitment to maintain the peg. Thus, sterilized intervention in this model is effective in the sense that it affects agents' *ex ante* assessment of the realignment probabilities. An increase in the amount of intervention decreases the probability of a realignment.

The model generates other interesting implications: First, the model predicts a relationship between intervention and the conditional variance[7], thus providing a rationale for empirically modeling the conditional variance as a function of intervention (e.g. within a GARCH framework). There are some empirical applications that report that intervention does not affect the mean but is volatility increasing.[8] In view of Isard's model, however, such a result does not necessarily imply that intervention is unsuccessful. In contrast, it could be perfectly compatible with the view that intervention exerts a stabilizing influence on exchange rates, but in a different sense, namely in influencing the realignment probabilities. By estimating a GARCH model, this latter effect cannot be detected (apart from the side effect on the conditional variance) thus casting doubts on the adequacy of the GARCH methodology in this context.

Second, the model predicts the presence of a peso problem. Each period, the govern-

[5]For example, if the authorities face speculative pressure to devalue and react by selling foreign currency, then a realignment will reduce the valuation gains that the government would have accrued in the absence of intervention.

[6]Notice, that similar to the signaling channel it is assumed that the intervention is publicly known. However, in contrast to the signaling channel, it is not assumed that central banks have more knowledge than private agents.

[7]The sign of this effect is ambiguous, depending on several model parameters.

[8]E.g. Dominguez (1998), Baillie & Osterberg (1997), for some EMS exchange rates analyzed in Brandner, Grech & Stix (2001).

ment faces an incentive to realign. Although actual realignments are infrequent and thus the probability is small, the presence of the peso problem influences the exchange rate distribution. And third, the realignment probabilities are serially correlated. This effect originates from the dependence of the realignment probabilities on output which itself is serially correlated. Finally, the model also predicts that realignment expectations depend on intervention in a nonlinear way.

As shown below, all of these implications can be exploited in choosing an empirical model to estimate the impact of intervention. In particular, the model's prediction that the focus of analysis should lie at the impact of intervention on the *ex ante* probability of a speculative attack will be tested.

An alternative approach is the one by Obstfeld (1994) who also models escape clause arrangements but additionally allows for the presence of self-fulfilling speculative attacks and thus multiple equilibria. In these models jumps from one equilibrium to another can cause speculative attacks without the presence of inconsistent fundamentals. The inability of models that build solely upon macroeconomic imbalances (e.g. Krugman 1979) to satisfactorily explain the EMS currency crises in 1992 and 1993 renders the multiple equilibrium approach very appealing for explaining currency crises (Rose & Svensson 1994). For example, it has been argued that uncertainties about the outcome of the ratification process of the Maastricht treaty and about the future prospects of monetary union lead to a shift in expectations resulting in attacks. In turn, these attacks forced governments to adopt policies (e.g. raising interest rates to prevent them) which were inconsistent with achieving the Maastricht criteria and made further maintenance of the peg very expensive in terms of domestic policy goals, like unemployment. This (newly created) inconsistency justified the speculative attack *ex post* although, in the absence of an attack, interest rates would not have been risen.

Similar to Isard's approach, Obstfeld (1994) assumes that the government has the option to abandon the peg if a large negative output shock hits the economy. By doing so, the policy maker can stimulate output (and thus employment) because of the presence of a short-run Phillips curve effect. Again, if the government devalues, it faces (reputational) fixed costs. Notice that unlike in Isard's (1995) model, there is no explicit role of intervention and thus valuation costs of foreign exchange reserves .

In equilibrium, Obstfeld's (1994) model predicts the presence of multiple equilibria. In this framework it is intuitively plausible that an increase in the fixed costs that the government has to incur when it abandons the peg will strengthen the commitment of the government and therefore makes devaluations less likely. However, due to the presence of multiple equilibria, this does not hold in general. As Flood & Marion (1999) and Sarno & Taylor (2001) discuss, in some equilibria the model can also predict a counter-intuitive result, namely that an increase in the costs will make devaluations more likely.

As argued by Sarno & Taylor (2001), a synthesis of Obstfeld's self-fulfilling attack story with Isard's approach of taking the valuation of foreign exchange reserves into account might have the potential of shedding new light on the question why central banks prefer to intervene secretly. According to this synthesis, the fixed costs in Obstfeld's (1994) model are interpreted as a transfer of wealth from central banks to speculators when a currency is successfully attacked—the valuation losses of the government are the profits of speculators. However as argued, Obstfeld's model predicts that an increase in these costs may result either in an increase or a decrease of the probability of a speculative attack. The sign of the effect depends on the particular equilibrium in which the economy has settled.

Within Obstfeld's (1994) model it is not possible to determine the particular equilibrium and when and why the economy jumps from one equilibrium to another. However, as remarked by Flood & Marion (1999), the Isard interpretation of intervention also has some merits in this context. If one is willing to interpret the valuation losses of the foreign exchange reserves as a transfer to speculators when the peg is abandoned, then it is likely that speculators prefer to settle on the particular equilibrium where speculators potentially earn most, or put differently, where attacks are most frequent. But this equilibrium is the one which produces the above mentioned counter-intuitive result where an increase in the costs—an increase in intervention—makes speculative attacks more likely. As Sarno & Taylor (2001) argue, it might be the presence of this perverse effect which explains why central banks prefer to intervene secretly.

As the previous discussion has shown, there are several potentially testable hypotheses that can be derived from this literature. First, Isard's (1995) model predicts that the *ex ante* probability of a speculative attack depends on intervention. And second, the above mentioned synthesis of Isard's (1995) and Obstfeld's (1994) model predicts that intervention might actually increase or decrease the probability of a speculative attack. If we find evidence in favor of a perverse effect of intervention, then this would provide support for Sarno & Taylor's (2001) explanation of central banks' tendency to intervene secretly.[9]

In the following section an empirical approach is proposed that allows to test these hypotheses.

2.2 The Markov Switching Model

Since the estimates will deal with target zone exchange rates it is necessary to select an empirical approach that is able to deal with the time series peculiarities (nonlinearities, fat tails, etc.) of target zone rates. One approach which seems to be able to capture these features is the Markov switching model (Engel & Hamilton 1990). Among others, Engel & Hakkio (1996), Martinez-Peria (2002) and Brandner et al. (2001) highlight the advantages of this model for EMS exchange rates and currency crises.[10]

In particular, the Markov Switching (MS) model can generate fat tailed and/or multimodal distributions as it assumes the presence of several distinct regimes. For example, if one regime is characterized by a conditional distribution with a large variance and the other by a conditional distribution with a small variance, then the unconditional distribution will have fatter tails than a normal distribution. Alternatively, if the two means of the distributions are different, then the unconditional distribution could have two modes. Second, both target zone models and speculative attack models predict a nonlinear re-

[9]Brandner, Grech & Stix (1998) report evidence that intervention activity by some EMS central banks increased the probability of shifting from a low to a high volatility regime. They, however, do not distinguish between secret and publicly recognized interventions and also analyze a different sample period.

[10]Like any techniques, the Markov switching model also has its drawbacks. For example, testing for the number of regimes is difficult.

lationship between target zone exchange rates and fundamentals.[11] The MS approach allows to model this. Third, the assumed serial correlation in the regime probabilities can generate volatility clustering which is frequently found in exchange rates (e.g. Engel & Hakkio 1996, De Vries 1992).[12] Finally, with the time-varying transition probability model it is possible to estimate the effectiveness of intervention conditional on a given regime.

Given these arguments, a simple MS model will be presented. In this model it is assumed that the exchange rate changes evolve according to,

$$\Delta s_t = \mu_{i,t} + \epsilon_{i,t} \tag{1}$$

where Δs_t is 100 times the log exchange rate return between time $t-1$ and time t, $\mu_{i,t}$ is a conditional mean and $\epsilon_{i,t}$ is an independently, identically and normally distributed error term with zero mean and variance σ_i^2. The subscript i for the conditional mean and error term indicates that the exchange rate return Δs_t is drawn from one of two distributions which are indexed by $Z_t = i$ $(i = 1, 2)$.[13] The assumption that the regime indicator variable Z_t evolves according to a first-order Markov chain implies that the conditional probability that the observation Δs_t has been drawn from distribution (regime) i at time t depends on the regime it has been drawn from at time $t-1$: $p^{ji} = P(Z_t = i \mid Z_{t-1} = j)$.

Although the regime indicator variable Z_t is not directly observable it is possible to draw some probabilistic inference about the regimes. For example, the *ex ante* probability $p_{i,t} = P(Z_t = i|\Omega_{t-1})$ measures the probability that the process is in regime i at time t conditional on the information set Ω_{t-1} which contains all available information up to time $t-1$.[14] Given the assumption of two regimes and about the distribution of the error term, the (composite) distribution of the exchange rate returns can be written as a mixture of

[11]For example, Krugman (1991) and Isard (1995).

[12]For example, see Hamilton & Susmel (1994).

[13]For the moment we assume that there are two regimes. Below, we provide evidence in favor of this assumption.

[14]The *ex ante* probability is the forecasted probability that the exchange rate will be in regime i at time t using information up to time $t-1$. The *ex post* probability is the updated probability using information up to time t.

two different normal distributions each weighted with probability $p_{i,t}$:

$$\Delta s_t \sim N(\mu_{i,t}, \sigma_i^2) \text{ with probability } p_{i,t} \quad i = 1, 2. \tag{2}$$

To complete the model, the conditional means for the two regimes are specified as follows: To allow for mean reversion in Δs_t we assume that the conditional mean in regime i is given by,

$$\mu_{i,t} = c_{i0} + c_{i1} \cdot D_{t-1} \quad , \tag{3}$$

where D_{t-1} denotes the percentage deviation from the D-mark central parity.[15] Notice that this formulation is very general in that it encompasses regimes with no mean reversion ($c_{i1} = 0$) and different degrees of mean reversion ($c_{i1} \neq c_{j1} \quad i, j = 1, 2$).

Although this time series model has been shown to have many advantageous features, it is unnecessarily restrictive in the sense that the transition probabilities p^{ji} are fixed in time. To relax this restriction, we follow the proposal of Diebold, Lee & Weinbach (1994) and model the transition probabilities as being time-varying. In this model the transition probabilities at time t are functions of a vector of predetermined variables:[16]

$$p_t^{ii} = f(\text{predetermined variables}_{t-1}), \quad i = 1, 2. \tag{4}$$

In particular, the transition probabilities are modeled as functions of intervention. By estimating the parameters of the functional specification $f(\cdot)$ one can test whether intervention is informationally relevant for the determination of regimes. The functional specification itself is of the logit type which guarantees that p_t^{ii} is in the range between zero and one,

$$p_t^{ii} = P(Z_t = i | Z_{t-1} = i; I_{t-1}) = \frac{e^{\beta_{i0} + \beta_{i1} \cdot I_{t-1}}}{1 + e^{\beta_{i0} + \beta_{i1} \cdot I_{t-1}}}, \quad i = 1, 2, \tag{5}$$

[15]A more detailed variable description is given in the Appendix on page 35.

[16]For a more detailed description of the model the reader is referred to Hamilton & Susmel (1994) and Diebold et al. (1994), for example.

where I_{t-1} is a measure of intervention specified below and β_{i0} and β_{i1} are coefficients to be estimated $(i = 1, 2)$.[17]

Given the complete model in equations (2), (3) and (5), the parameters which need to be estimated are $c_{i0}, c_{i1}, \sigma_i^2, \beta_{i0}$ and β_{i1} $(i = 1, 2)$. This is done by numerically maximizing the log-likelihood function as described in Diebold et al. (1994).[18] Whether intervention is informationally relevant in the two regime model can be tested by means of a likelihood ratio test of the null hypothesis $\mathbf{H}_0 : \beta_{11} = 0, \beta_{21} = 0$ against the alternative hypothesis $\mathbf{H}_\mathrm{A} : \beta_{11} \neq 0, \beta_{21} \neq 0$. The resulting test statistics is chi-square distributed with two degrees of freedom. Alternatively, it is also possible to test for the significance of the individual parameters β_{i1} by means of Wald tests.[19]

Apart from the advantages listed above there are two more arguments which are directly related to the economic models discussed above and which make the time varying Markov Switching model particularly appealing in the context of speculative attack models. The first argument rests on the model of Isard (1995), the second on the model of Obstfeld (1994).

As discussed above, Isard's (1995) model predicts that the *ex ante* probability of a speculative attack is a function of the amount of intervention. This is nicely incorporated in the Markov Switching model above where the *ex ante* probability $P(Z_t = i|\Omega_{t-1})$ is given by

$$\begin{aligned} P(Z_t = i|\Omega_{t-1}) =& P(Z_t = i|Z_{t-1} = i; I_{t-1}) \cdot P(Z_{t-1} = i|\Omega_{t-1}) \\ &+ P(Z_t = i|Z_{t-1} = j; I_{t-1}) \cdot P(Z_{t-1} = j|\Omega_{t-1}) \ , \end{aligned}$$

where $P(Z_{t-1} = k|\Omega_{t-1})$ is the time $t-1$ *ex post* probability and $P(Z_t = i|Z_{t-1} = k; I_{t-1})$ (for $k = i, j$) are the time varying transition probabilities as functions of intervention I_{t-1}. Thus, if for example regime i is a speculative attack regime then clearly the *ex ante*

[17] Of course, if $\beta_{i1} = 0$ the model collapses to the fixed transition probability case $(i = 1, 2)$.

[18] Further estimation details can be found in a supplement which is available from the author.

[19] Filardo (1998) derives conditions for the time-varying transition probability model such that the estimated maximum likelihood parameters have their desired asymptotic properties.

probability that a "speculative attack" will happen next period, $P(Z_t = i|\Omega_{t-1})$, depends on intervention. More specific, if regime one is a calm regime and regime two a speculative attack regime, then Isard's (1995) model would predict that the *ex ante* probability of the speculative attack regime decreases with the amount of intervention ($\frac{\delta P(Z_t=2|\Omega_{t-1})}{\delta I_{t-1}} < 0$). A sufficient condition for this to be the case is that the conditional probability that the process stays in the calm regime once it is there increases ($\beta_{11} > 0$) and the corresponding conditional probability of the speculative attack regime decreases ($\beta_{21} < 0$).[20]

Although obvious, it should be mentioned that two additional features of Isard's (1995) theoretical model are incorporated in the MS model: First, due to the presence of two regimes it takes account of the peso problem and second, due to the Markov chain assumption it can generate serial correlation in the *ex ante* probabilities.

The second theoretical argument rests on the escape clause model presented above which states that a rise in the cost of exciting the escape clause can have two opposing effects depending on the respective equilibrium: in the "normal" equilibrium the probability of a speculative attack decreases with the amount of intervention and in the "perverse" equilibrium the same probability will increase. Given estimates of β_{11} and β_{21}, it is possible to test which equilibrium prevailed: Again, if regime two is the speculative attack regime, then $\beta_{11} < 0$ jointly with $\beta_{21} > 0$ implies, that an increase in the amount of intervention increases the *ex ante* probability of a speculative attack ($\frac{\delta P(Z_t=2|\Omega_{t-1})}{\delta I_{t-1}} > 0$)—thus this parameter constellation would characterize a perverse effect of intervention. If $\beta_{11} < 0$ and $\beta_{21} = 0$ then only those interventions undertaken in the calm regime increase the probability of going to the speculative attack regime while those undertaken in the speculative attack regime do not matter.

As discussed, the perverse effect of intervention rests on the fact that interventions are

[20] In general, the sign of the derivative of the transition probability function $P(Z_t = i|Z_{t-1} = i; I_{t-1})$ with respect to intervention corresponds to the sign of β_{i1} (compare with eq. 5): If $\beta_{11} < 0$, then this implies that $\delta p_t^{11}/\delta I_{t-1} < 0$. If $\beta_{21} > 0$, then this implies that $\delta p_t^{22}/\delta I_{t-1} > 0$. This particular parameter constellation would thus imply that (a) the probability that the process stays in regime 1 once it is there decreases, and (b) the probability that the process stays in regime 2 once it is there increases with the amount of intervention. *Ex ante*, this makes it more probable that the process will be in regime 2 next period.

public. Therefore, we should only observe significant $\beta_{11} < 0$ and $\beta_{21} > 0$ for those interventions that were publicly recognized but not for secret interventions. In turn, significant $\beta_{11} > 0$ and $\beta_{21} < 0$ would indicate that an increase in intervention makes a speculative attack less likely.

3 Data Description and Stylized Facts

The data used in this paper comprise the French franc and Spanish peseta exchange rates vis-à-vis the Deutsche mark as well as data about interventions undertaken by de Banco de España, the Banque de France and the Deutsche Bundesbank in the period from June 1, 1992 to July 30, 1993.[21] Thus, the sample covers the very important period of speculative attacks and pressure during 1992 and 1993.[22]

Concerning the exchange rate and sample choice some remarks are necessary. To test the hypotheses discussed above, it is necessary to have a sample of exchange rates that were exposed to speculative pressure. This sample should be as long as possible and importantly, it should also contain enough intervention events. Such a data set is not easy to find as typically, intervention data are not publicly available. Moreover, to separate the effects of public and secret interventions it is necessary that the data set comprises enough interventions of each category. These reasons limit the choice to those exchange rates analyzed in this paper.[23] While highly desirable an analysis of more exchange rates and different sample periods is not possible due to data constraints.

[21]The beginning of this period is given exogenously due to data constraints and the end date is chosen because it represents the end of the narrow band EMS.

[22]For a summary of the events in the EMS during 1992 and 1993, the reader is referred to Eichengreen (2000).

[23]For example, the vast majority of interventions in the EMS after 1993 were secret leaving only a few public interventions which precludes the use of the post 1993 period. In the pre 1993 period considered here, sufficiently many intervention events are only available for a few exchange rates. Unfortunately, data on UK sterling and Italian lira are not available to the author.

3.1 Exchange Rate Data and Stylized Facts

The exchange rate data are derived from Bank for International Settlement USD exchange rate series, laid down at the daily concertation procedure of central banks at 14:15.[24] The DEM cross rates are calculated by assuming that the no-triangular-arbitrage condition holds. Exchange rates (S_t) are expressed in terms of DEM per 100 units of local currency.[25] The exchange rate returns (Δs_t) are calculated as 100 times the log difference of the exchange rate.[26]

Table 1 presents some descriptive statistics for the exchange rate returns. The results strongly emphasize the importance of well known stylized facts of target zone exchange rates (De Vries 1992, Brandner et al. 2001): excess kurtosis and skewness (for the peseta). As expected, the Jarque-Bera test for normality is overwhelmingly rejected for both the peseta and the franc. The results for the $Q^2(20)$ test strongly indicate the presence of volatility clusters.

The daily percentage exchange rate changes are depicted in Figure 1 which shows the typical pattern of volatility clusters as well as the presence of considerable within-the-band changes. Typically, such large changes occur in speculative attack periods. In general, the fluctuations are much larger for the peseta than for the French franc which can be attributed to the fact that the peseta was in a 6% band whereas the franc was in the narrow 2.25% band. Also, the peseta was realigned three times within the sample.[27]

3.2 Intervention Data

The intervention data used in this paper are daily cumulated (net) DEM-intervention volumes by the Banco de España and the Banque de France. In particular, as both currencies

[24] The following abbreviation will be used: ERM for Exchange Rate Mechanism, DEM for Deutsche mark, ESP for Spanish peseta and FRF for French franc. "Peseta" and "franc" refer to the ESP/DEM and FRF/DEM exchange rate, respectively.

[25] An appreciation means that $S_t > S_{t-1}$.

[26] The series used in this paper are described in more detail in the Appendix (page 35).

[27] The largest negative change of the peseta is due to the realignment on September 17, 1992, the extreme positive change of 4.44% to the imposition of capital control on September 23, 1992.

came under speculative attack and weakened against the mark, we concentrate only on the DEM-amounts spent by the respective central banks, leaving purchases aside.[28] The daily intervention figures are collected from concertation protocols. Each day, all EMS and a few other central banks communicate the amounts of DEM-purchases and DEM-sales in four concertation rounds. The first round takes place at 9:30 and the last round at 16:00. The intervention data is cumulated over a time period of 24 hours, starting from 16:00 previous day until 16:00 today. Intervention undertaken after 16:00 are reported at the first concertation round next day at 9:30 and are therefore included in next day's intervention figure.[29]

Since the ERM I was a multilateral target zone, the fluctuation band was not only defined vis-à-vis the German mark but also vis-à-vis each other participating currency. Thus, intervention obligations could have arisen between two partner central banks without necessarily involving DEM operations. To avoid that non-DEM-interventions affect the results, it is necessary to select only currencies for which by far the most intervention activity was conducted in DEM to influence the bilateral mark exchange rate. An analysis of the intervention data in combination with careful reading of news reports shows that this is the case for the peseta and the franc.

We also want to analyze the influence of Bundesbank interventions on the exchange rates which is not directly possible because we do not have access to the actual Bundesbank intervention amounts.[30] Hence, a constructed series is used which relies on information on the Bundesbank's system wide intervention data and on news reports. It is well known that the Bundesbank did not intervene intra-marginally, except in very rare circumstances—for example in the fight for the franc in September 1992. In these cases and additionally in

[28]As DEM-purchases were mainly undertaken to replenish reserves and not to influence exchange rates, DEM-sales were the relevant operations.

[29]Readers will notice that exchange rates are recorded at 14:15 while intervention operations are recorded at 16:00. Unfortunately, this timing discrepancy is unavoidable. However, a careful analysis of news reports allows the conclusion that the vast majority of interventions were undertaken before 14:00. This fact attenuates this inconsistency.

[30]In the data set used, Bundesbank intervention figures are only available as aggregate figures for all EMS currencies. Therefore, it is not possible to derive the intervention volume for individual bilateral exchange rates from this source of information.

the cases when the Bundesbank was obliged to intervene, public awareness was given.[31]. Thus, the constructed series is a dummy series taking a value of one if the Bundesbank was reported to be on the market and where we observe Bundesbank system wide intervention.

To be able to test whether the effect of intervention differs across reported and secret interventions the database of the Reuters News service and the Financial Times are searched on a daily basis for news about central bank activities.[32] Usually both sources are well informed and report comprehensively about interventions. Both archives together should therefore yield a reliable and informative picture about financial markets' awareness of intervention activity. Given this information, a "news" variable is constructed which, on a given day, takes a value of one if there was a report about intervention activities and zero else. Whether an intervention was known on financial markets or not was judged according to a rather conservative criterion: whenever on a given day there was any indication or any suspicion of central bank intervention, the dummy variable is set to one.[33] The exchange rates and the days when news about interventions occurred are shown in Figure 2 and 3.

In Table 2 the constructed intervention news series is compared with the actual intervention data.[34] DEM-sales by the Banco de España and the Banque de France were reported on 50 and 38 days, respectively. On 18 days the Bundesbank intervened in support for the French franc but never for the peseta. In comparison, the actual intervention data show that the true numbers of intervention days are 73 and 58 for the Banque de

[31]For example: "The Bundesbank intervened to support a currency within its margins for the first time since the ERM was founded." (Financial Times, 'UK: Foreign Exchanges - Battle to save the French franc', September 24, 1992, page 35). On days when the Bundesbank intervened intra-marginally, it published its intervention activity anyway. For example, on September 23, 1992, the central banks and finance ministries of France and Germany issued a joint statement declaring their intervention policy to the public.

[32]The archive used was Reuters Business Briefing. The sources Reuters News and Financial Times were searched for the topics "Central Bank Intervention" or "Foreign-Exchange News" or "Money and Forex Markets".

[33]This includes cases when some traders suspected intervention while others were not sure. Notice that we do not separate between interventions that are announced, confirmed or denied by central banks. This would be virtually impossible because it could be the case that a central bank intended to announce its intervention activity but then abstained from doing so because the activity got public anyway. Therefore, the minimum requirement for an intervention to be counted as "publicly known" is a rumor.

[34]Since the intervention figures used in this paper are not publicly available, we are only able to present some aggregate statistics such that individual daily operations are not revealed.

France and the Banco de España, respectively. Thus, 34 interventions by the Banco de España and 26 by the Banque de France remained secret. By and large, this analysis shows that news reports about interventions do not match the actual intervention events accurately: there are days in the sample when an intervention was reported that did not take place and days with (explicit) statements of no observed intervention when in fact there was some intervention activity.

Given these data, it is analyzed next how interventions—and in particular, how secret and public interventions—affected the DEM/ESP and the DEM/FRF exchange rates.

4 Results from the Markov Switching Model

The estimation results for the peseta and the franc are summarized in Table 3 and 4. The estimation strategy was to find a preferred fixed transition probability model whose results are presented in column I.[35]

For the peseta one regime is found which is characterized by a significant negative drift (minus 0.03% daily change) and low variance and another regime with a more pronounced negative drift (minus 0.33%) and higher variance.[36] The transition probability parameter β_{10} and β_{20} are 2.38 and 0.85 and both significant at a 5% level. These parameter estimates imply that the exchange rate stayed in the first regime once it was there with a probability of 0.92 and in the second regime with a probability of 0.70. Table 3 also presents the results of Ljung-Box $Q(20)$ and $Q^2(20)$ tests. Both test statistics show that the null of no serial correlation in the standardized and squared standardized residuals cannot be rejected at conventional levels of significance.

For the French franc the results of the specification search yield a model in which the first regime is characterized by a driftless random walk (as c_{10} is not significantly different

[35]Given the relatively low number of observations, estimating a large general model is not possible. Therefore, a simple-to-general specification search was chosen. The details of this specification search are not shown but can be obtained from the author. In general, all estimation results are obtained via numerical optimization with the GAUSS optmum package.

[36]The variance in the second regime is about 60 times higher than the variance in the first regime.

from zero) and the second regime is characterized by mean reversion. Comparable to the peseta, the variance of the second regime is found to be higher than the variance of the first regime (seven times). In absolute terms the variance in this high variance regime is estimated to be much smaller for the franc than for the peseta (which was expected because the franc was not realigned during the sample). Also, the two models differ in the sense that there is mean reversion for the French franc but not for the peseta. Most likely this reflects the presence of realignments for the peseta while the franc's central parity did not change. The parameters of the transition probability function, again both significant at a 5% level, imply transition probabilities of 0.95 (regime one) and 0.91 (regime two), respectively. Hence, the expected duration of the high variance regime was higher for the franc (11.4 days) than for the peseta (3.3 days). Also, for the franc there is no evidence of serial correlation in the standardized and squared standardized residuals.

The *ex post* probabilities of the second regime for both exchange rates are plotted in Figure 4. For the French franc, the first hike in the probability of the second regime occurred when the British Pound repeatedly came under heavy pressure putting the EMS under strains.[37] The shaded area in Figure 4 shows the period from the Lira realignment (September 14) until the appearance of press reports about the easing of the tension on the franc (September 29).[38] As can be seen, during most of this period, the probability that the franc was in the second regime was very high. The next increase of the high variance regime probability occurred around November 20, where "speculation that another realignment of parities within the ERM ... [was] waiting in the wings intensified ..."[39]. In a similar manner, also the other hikes can be attributed to periods of tensions and speculative attacks.

For the peseta, in general the picture is the same: increases in the probability of regime two can be attributed to periods of speculative pressure. In Figure 4 the realignments

[37] Aug. 24, 1992

[38] "The French franc strengthened against the D-mark, indicating that devaluation of the currency looks unlikely", Financial Times, "UK: Sterling hits new low as franc rallies", Sept. 29, 1992, page 1.

[39] Reuters News Service, "France: ERM realignment speculation grows", Nov. 20, 1992.

and the imposition of capital controls are marked by bold vertical lines.[40] Typically, the probability of regime two is very high already before these events. Therefore, given this evidence and the pattern of regime probabilities, regime one is interpreted as the "calm" and regime two as the "speculative attack" regime.[41]

Next, based on the preferred models, the intervention variables are included in the transition probabilities. First, the results for the dummy variable "news" are shown in column II of Table 3 and 4. Again, if interventions cause an increase in the probability of a speculative attack then we should find $\beta_{11} < 0$ and/or $\beta_{21} > 0$.[42] For both currencies the estimated effect of news given that the exchange rates were in the "calm" regime (β_{11}) is negative, although statistically significant so only for the peseta. Additionally, for the peseta the likelihood ratio test of $\beta_{11} = 0$ and $\beta_{21} = 0$ is rejected at a 10% level. In contrast, both β_{21} coefficients are individually insignificantly different from zero. This result implies, that the probability that the DEM/ESP exchange rate stayed in the calm regime was reduced from 0.94 to 0.75 whenever news about intervention by the Banco de España appeared. Overall, there is some evidence that intervention news increased the probability of a speculative attack for the peseta but not so for the French franc.

As argued, the scope of the "news" dummy variable is conservative in the sense that a rumor on financial markets gets the same weight as a massive intervention implying that traders would react to a significant intervention event the same as to a negligible rumor. If traders are more cognizant about intervention and observe both the fact that central banks intervene as well as the approximate amount, then the "news" variable might be a bad proxy for the information set of traders.[43] To take this into account, the Markov Switching model is estimated with the intervention amount entering the transition probabilities. These results are shown in column III. Here, the likelihood ratio test of zero coefficients—

[40] The first, third and fourth line shows the realignments, the second the imposition of capital controls.

[41] An alternative labeling of regimes could be that regime one is interpreted as the "low volatility" and regime two as the "high volatility" regime.

[42] Compare with footnote 20 on page 17.

[43] Typically, traders have a good feeling for the approximate size of intervention. Subjective categorizations as "massive" or "moderate" can be found quite often in the press. In addition, if amounts are reported, traders' estimates of the amounts often are close to the actual figures.

$\beta_{11} = 0$ and $\beta_{21} = 0$—is rejected for both currencies. Furthermore, for the peseta, β_{21} is significantly different from zero and positive—thus indicating that it is more likely that the peseta stayed in the turbulent regime if DEM-sales by the Banco de España occurred. For the calm regime, the parameter (β_{11}) is not significant. For the French franc β_{11} is significantly negative. Thus, if the Banque de France sold D-mark in the calm regime, the probability of staying in the calm regime decreased.

Next, DEM-sales are split into those that were known to the public and those that stayed secret (columns IV and V). For the publicly recognized interventions, β_{11} is negative and (individually) significant for the peseta (at a 5% level) and for the franc (at a 10% level). In addition, β_{21} is weakly significant and positive for the peseta. Furthermore, the likelihood ratio test rejects the null of jointly zero β_{11} and β_{21} coefficients at the 5% level.

As discussed, during the 1992-1993 EMS crises the Bundesbank intervened—even intra-marginally—in support for the French franc. In particular, the September 1992 speculative attack against the franc was, according to many observers, successfully fend off because of this Bundesbank support. Because of the particular role of the Bundesbank in the EMS, it is therefore likely that Banque de France interventions that were undertaken without Bundesbank support had a different impact than joint interventions. To account for this, column V of Table 4 reports the results for interventions that were undertaken by the Banque de France alone. For these interventions, we find that β_{11} is negative and significant on a 5% level.

The estimated parameters for secret intervention are not significant for the franc. For the peseta β_{11} is significant at a 10% level and positive, implying that secret interventions by the Banco de España increased the probability of staying in the calm regime.[44]

It could be the case that the results are caused by large interventions and not by the fact that the interventions were public. This possibility is tested by defining a dummy variable for public interventions thus treating small and large public interventions the same. If the

[44] However, this result must be interpreted with some care as the estimate of β_{21} is very imprecise. This is likely to be due to the fact that the number of secret interventions in the "speculative attack" regime is small.

effect of this dummy variable is estimated, then the results for the FRF/DEM rate are qualitatively the same.[45] For the peseta, this model did not converge.

In sum, the point estimates for both exchange rates show that public interventions without Bundesbank support undertaken in the "calm" regime increased the probability that the exchange rate went into the "speculative attack" regime.

5 Is there an Impact of Intervention on Realignment Expectations?

So far, the analysis rested on empirical models of changes of the daily spot exchange rate. It was estimated how intervention affected the probability of next day's exchange rate regime. However, intervention can also have a different impact by affecting the expectations of market participants. In principle, option prices could be used to derive the distribution of the expected exchange rate. As we do not have access to option data from this early period we instead calculate devaluation expectations along the line of Svensson (1993) and Rose & Svensson (1994) for both the peseta and the franc exchange rate to analyze this channel.

Rose & Svensson (1994), studying EMS currencies until 1993, find no economically meaningful relation between realignment expectations and macroeconomic variables. In contrast, Koedijk et al. (1995), analyzing intervention data by the Banque Nationale de Belgique, find that intervention at the weak edge of the band signals weakness and may therefore "even help to undermine the currency" (ibid., p.508). The sample as well as the exchange rate in Koedijk et al. (1995) is different than ours. Therefore, it is of interest first, whether this result can be confirmed for the franc and the peseta during 1992 and 1993 and second whether the use of actual intervention amounts as opposed to intervention dummies in Koedijk et al. (1995) gives similar results.

[45] Moreover, the statistical fit improves as β_{11} is significantly different from zero on a 5% level for public DEM-sales and DEM-sales without Bundesbank support. This result is not shown but can be obtained from the author.

Svensson (1993) and Rose & Svensson (1994) propose to estimate exchange rate credibility by either the plain or an adjusted measure of the interest rate differential where, according to Svensson (1993), the latter yields more precise estimates than the former for short horizons. In particular, Svensson shows that the expected rate of realignment can be decomposed into the expected rate of depreciation minus the expected rate of depreciation within the band. By assuming uncovered interest parity, the expected rate of depreciation can be measured by the differential of the domestic and the German interest rate. The expected rate of depreciation within the band can be estimated.[46] As is well known, the such obtained estimated expected rate of realignment is the product of the expected realignment size and the expected realignment frequency.[47]

For each day in the sample we calculate the expected rate of realignment for a 1 month maturity.[48] Denoting this expected rate of realignment by g_t we then run the following regression,

$$g_t = c + \sum_{j=1}^{k} \gamma_j \cdot g_{t-j} + \alpha \cdot I_{t-1} + \beta \cdot (i_{t-1}^{3m} - i_{t-1}^{1m}) + \sqrt{h_t} \cdot \epsilon_t \ ,$$

where ϵ_t is assumed to be independently, identically and normally distributed with zero mean and unit variance. I_{t-1} represents various measures of intervention. Notice that we use lagged intervention to avoid problems with simultaneity.

Certainly, one of the main purposes of intervention is to restore credibility in periods when markets are turbulent. If interventions are effective in this sense we should be able to find $\alpha < 0$.[49] However, given our previous results and the findings in Koedijk et al. (1995) it cannot be excluded *a priori* that the expected rate of realignment increases if

[46]For further details, the reader is referred to Svensson (1993).

[47]If it is assumed that the expected realignment size is a constant then the expected realignment rate is proportional to the expected realignment probability.

[48]The estimates are based on the one month differential of Euro bond interest rates as well as estimates obtained from applying equation 3 of Rose & Svensson (1994) to our data. In case of the peseta we adjust for the realignments as described in Svensson (1993). The estimates and graphs of the expected realignment rates can be obtained from the author.

[49]$g_t > 0$ represents expectations about a weakening of the home currency. $g_t < 0$ denotes revaluation expectations.

agents observe central banks intervening.

A positive $\alpha > 0$ could also arise if interventions are not (fully) sterilized. Then, purchases of local currency (sales of DEM) will have an increasing effect on domestic interest rates thereby increasing the interest rate differential vis-à-vis Germany. A positive coefficient on lagged intervention could just reflect this effect because our measure of the expected rate of realignment is positively correlated with the interest rate differential. Therefore, to isolate the unsterilized part of intervention we follow Koedijk et al. (1995) and include the slope of the yield curve as measured by the difference between the three month and the one month interest rate (i^{3m} and i^{1m}).

The point estimates of α are summarized in Table 5. For the peseta we find that an equation estimated by OLS with two lags of the dependent variable removes serial correlation in the residuals. Also there is no evidence of remaining ARCH effects. For the DEM/FRF exchange rate there are ARCH effects remaining, thus an EGARCH model is estimated for this currency.[50]

As shown in Table 5, the point estimates of α for lagged DEM-sales (line I) are positive and significant on a 5% level for the peseta and on a 10% level for the franc. Thus, this result is in line with the result reported in Koedijk et al. (1995). Decomposing interventions into publicly known and secret interventions (line II), shows that all estimated parameters for public interventions are significantly positive on a 5% level. In contrast, secret interventions are not found to be significantly different from zero. In the case of the franc, interventions may also influence the conditional variance of the expected realignment rate. These coefficients may be interpreted as the marginal effect of intervention on "uncertainty" about realignment expectations. As can be seen, only the coefficient for public intervention is significantly positive. Using only those interventions that were conducted by the Banque de France alone (line III) gives a different picture: there is no significant influence on the level but a significant and positive effect on the volatility of the expected realignment rate. Therefore, it seems that the positive effect of overall DEM-sales can be

[50]For the peseta two lags and for the franc one lag of the dependent variable is included.

traced to the effect of publicly known interventions.

6 Discussion of Results

The results from the Markov switching model lend statistically significant support to the view that interventions are informationally relevant for the determination of exchange rate regimes. News about intervention activity, irrespective of whether true or false, increased the probability that the DEM/ESP exchange rate jumped from the calm to the speculative attack regime. Significant results are also found for the amount of DEM-sales: According to likelihood ratio tests, DEM-sales influenced the regime probabilities significantly. Splitting interventions according to whether they are known to the market or not, shows that the relevance of DEM-sales for the regime probabilities can be attributed to publicly recognized but not to secret interventions.

As the individual parameter signs show, public (without Bundesbank support) interventions undertaken in the "calm" regime increased the probability that the exchange rate went into the "speculative attack" regime. This effect is significant on a 5% level for both exchange rates analyzed. In turn, interventions in the "speculative attack" regime are not found to be different from zero apart from one case where the probability of staying in the turbulent regime increases with intervention.

The estimates of the effects of intervention on market expectations yield a somewhat similar picture as public but not secret interventions are found to be positively correlated with the expected rate of realignment.

In terms of the theoretical models discussed above, none of these results do support the view that interventions decreased the probability of a speculative attack or the expected rate of realignment. Furthermore, the results from splitting interventions into publicly known and secret interventions shows that intervention has to some extent the "opposite" effect of what one might expect. The result thus show that the idea of a perverse effect of intervention cannot be rejected on the basis of the previous results.

How robust are these results? One objection raised against this interpretation of the results is that, different to the implicit assumption made above, interventions could be caused by market turbulences or by increases in the expected realignment rates and not the other way around. To test for the possibility of reversed causality, Tobit reaction functions are estimated. In this approach the amount of intervention is modeled as,

$$I_t^* = c + \alpha \cdot D_{t-1} + \beta \cdot (h_{t-1} - h_{t-1}^*) + \gamma \cdot g_{t-1} + \epsilon_t \tag{6}$$

where I_t^* denotes the censored intervention variable for which $I_t = 0$ if $I_t^* \leq 0$ and $I_t = I_t^*$ if $I_t^* > 0$. Intervention is thus modeled as depending on lagged deviations from the central parity (D_{t-1}) and deviations of the conditional variance from a target level ($h_{t-1} - h_{t-1}^*$), where the latter is defined as the 20-day moving average of the conditional variance.[51] Furthermore, the amount of intervention depends on the expected rate of realignment lagged one day (g_{t-1}).[52]

The estimation results are summarized in Table 6. For DEM-sales, the deviation from the central parity is significant at a 5% level for both exchange rates. As expected, negative deviations from the central DEM-parity had a positive impact on the amount of intervention. The deviation from the target rate of volatility had a positive and significant impact for the peseta for DEM-sales and for public DEM-sales. In contrast, secret DEM-sales by the Banco de España did not seem to be caused by deviations from target volatility. For the DEM/FRF exchange rate, deviation from target volatility did not cause intervention—except for those interventions that are undertaken by the Banque de France ("Public Sales w/o BUBA") alone.

[51]The conditional variance is derived from two EGARCH models that were estimated. The EGARCH models are specified as follows: For the peseta the mean equation contains the one month Eurobond interest rate differential vis-à-vis Germany, a realignment dummy and a dummy for the imposition of capital controls. The conditional variance equation contains the absolute deviation from the central parity. For the franc, the conditional mean contains the deviation from the central parity. The use of a window of 20 days for the conditional variance is arbitrary. The use of 10 or 5 day windows does not change the general conclusion from the results. All models and results are described in a supplement to this paper that is available from the author.

[52]In this formulation it is implicitly assumed that any deviation from perfect credibility ($g = 0$) affects intervention activity. We also experimented with a formulation where only realignment expectation (and not revaluation expectations) cause DEM-sales. The results are qualitatively the same in both specification.

The results of the Markov switching model imply a significant effect of public (peseta) and public-without-Bundesbank (franc) interventions conditional on taking place in the "calm" regime (regime 1). As can be seen from Table 6, these interventions do not seem to be caused by deviations from target volatility.

Significant parameter estimates of γ would provide evidence in favor of the reversed causality view in the case of the regression of the expected realignment rate on lagged interventions. The parameter estimates in Table 6 show that the possibility of reversed causality cannot be excluded for interventions by the Banco de España which were significantly caused by increases in g_{t-1}. However, for interventions undertaken by the Banque de France, no evidence of reversed causality is found.

On balance, the reaction functions thus show that the relevant Markov switching results for the franc and the peseta do not seem to be plagued by reversed causality problems. Concerning the results for the expected realignment rate, there is evidence of reversed causality in the case of the peseta.

7 Conclusion

In this paper the effectiveness of interventions undertaken by de Banco de España and the Banque de France during the period of speculative pressure in 1992 and 1993 is analyzed. In particular, we estimate a Markov Switching model of the daily change in the spot exchange rate in which the amount of DEM-sales influences the transition probabilities between regimes. It is argued that this model is in line with various implications of theoretical speculative attack models. Additionally, it is analyzed how interventions affected expectations about realignments that were estimated as proposed in Svensson (1993) and Rose & Svensson (1994). In both empirical approaches we study whether the effect of publicly known interventions differs from those that remained secret. This allows to test for the presence of a perverse effect of publicly known intervention as recently conjectured in Sarno & Taylor (2001).

The estimation results can be summarized as follows: In the Markov switching model we find evidence of the presence of two regimes which are interpreted as "calm" and "speculative attack" regimes. Interventions are informationally relevant for the determination of exchange rate regimes as the amount of DEM-sales significantly influenced the probabilities of switching between these regimes. For both exchange rates, there is evidence that public interventions (without Bundesbank support) in the "calm" regime significantly increased the probability that the exchange rates went to the "speculative attack" regime.

In regressions of the expected realignment rate on lagged interventions it is found for both exchange rates that the parameters of public interventions are significantly positive. Thus, public interventions seem to be associated with increases in the level and volatility of the expected rate of realignment. Secret interventions do not seem to have a significant effect.

In general, the results obtained in this paper do not correspond with the view that central bank interventions were able to calm markets or market expectations. The estimation results suggest the opposite, namely that publicly known interventions increased the volatility of the daily spot exchange rate and the expected rate of realignment. Therefore, up to the extent that our results represent a causal relationship, this finding provides support for the presence of a perverse effect of intervention.

Due to the peculiarities of the analyzed sample one should not push policy conclusion too far. However, given the importance attached to intervention in target zone arrangements like the EMS II, the results of this paper suggest that the role of intervention in periods of speculative pressure deserves further attention. Clearly, an analysis of more exchange rates and longer periods of intervention could shed more light on this issue.

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A Appendix

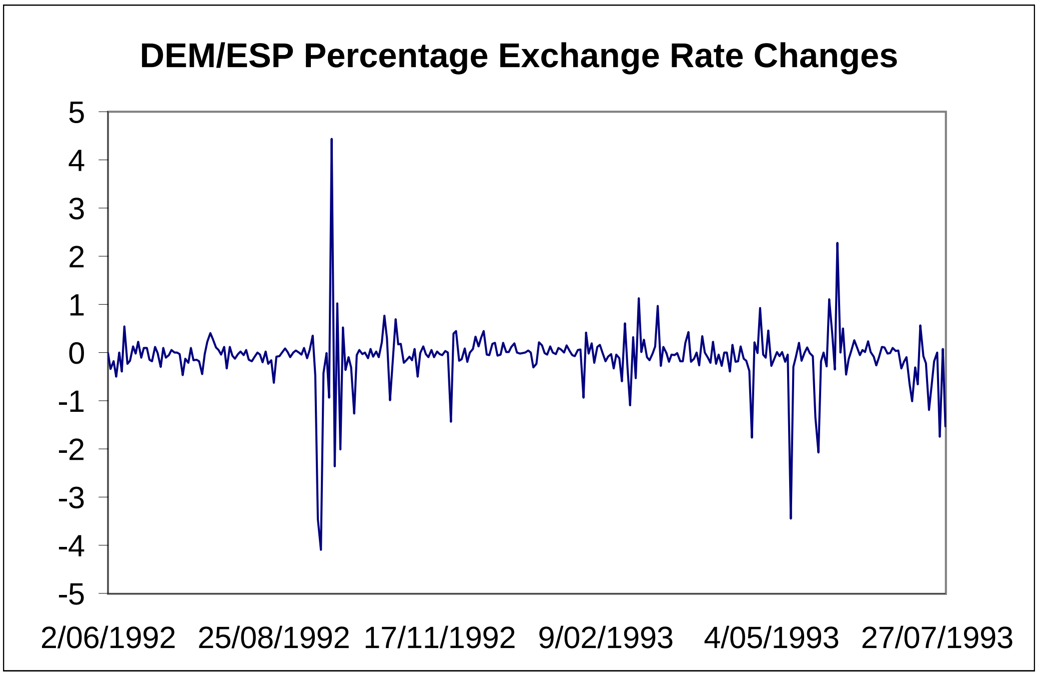
Data Description

- Δs_t ... The dependent variable is 100 times the log difference of 100 units of local currency expressed in Deutsche mark (DEM). An appreciation of the local currency corresponds to a positive change. The exchange rate series is taken from the BIS database (USD exchange rates). The USD rates are converted to bilateral rates by assuming that the no-triangular-arbitrage condition holds. The exchange rates are recorded at 14:15 European Central Time.

- I_t ... the amount of sales by the local central bank expressed in DEM as reported in the daily concertation rounds (source: own calculation). If on the same day a central bank sold DEM and purchased DEM, then we take the net (cumulated) amount of interventions. Typically, this is not restrictive as central banks barely change the intervention direction within a day. Notice, that interventions are recorded from 16:00 previous day until 16:00 today which does not exactly match the exchange rate definition. However, we do not expect this to influence the results because, reportedly, interventions barely occurred between 14:15 and 16:00.

- $D_t = \log S_t - \log \text{Central Parity}_t$ ... the log deviation of the exchange rate from the central parity. The variable used in the tables is divided by the width of the permissible fluctuation band (0.0225 for the franc and 0.06 for the peseta).

- $(h_{t-1} - h^*_{t-1})$ in the reaction functions measures the deviation of the conditional variance from a target level, where the latter is defined as the 20-day moving average of the conditional variance. The conditional variance itself is derived from EGARCH models. The EGARCH models and the results are described in a supplement to this paper that is available from the author.



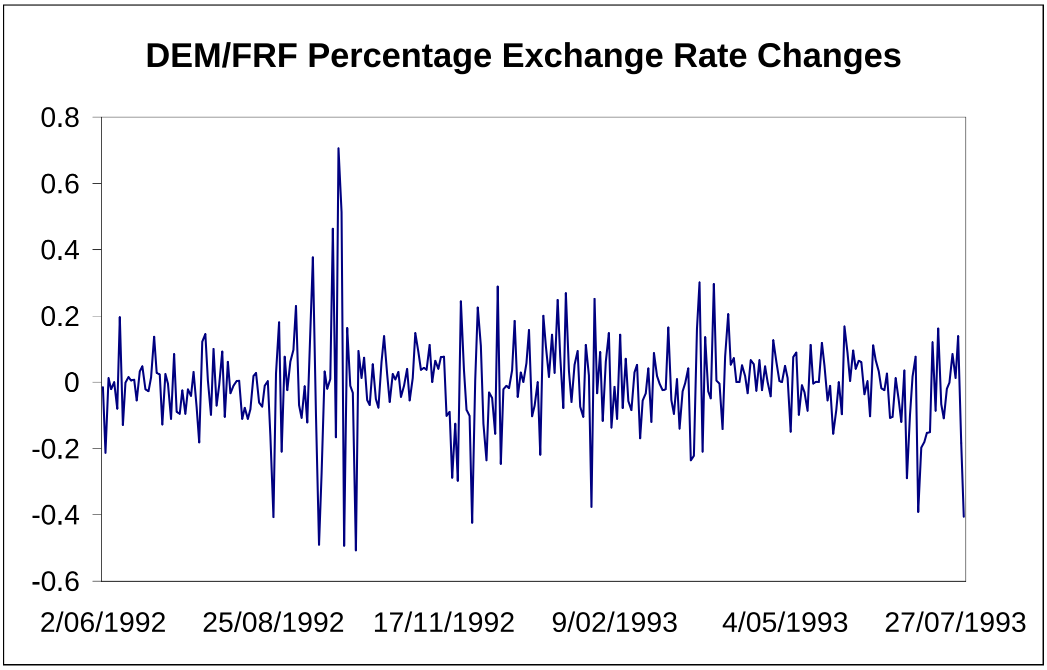
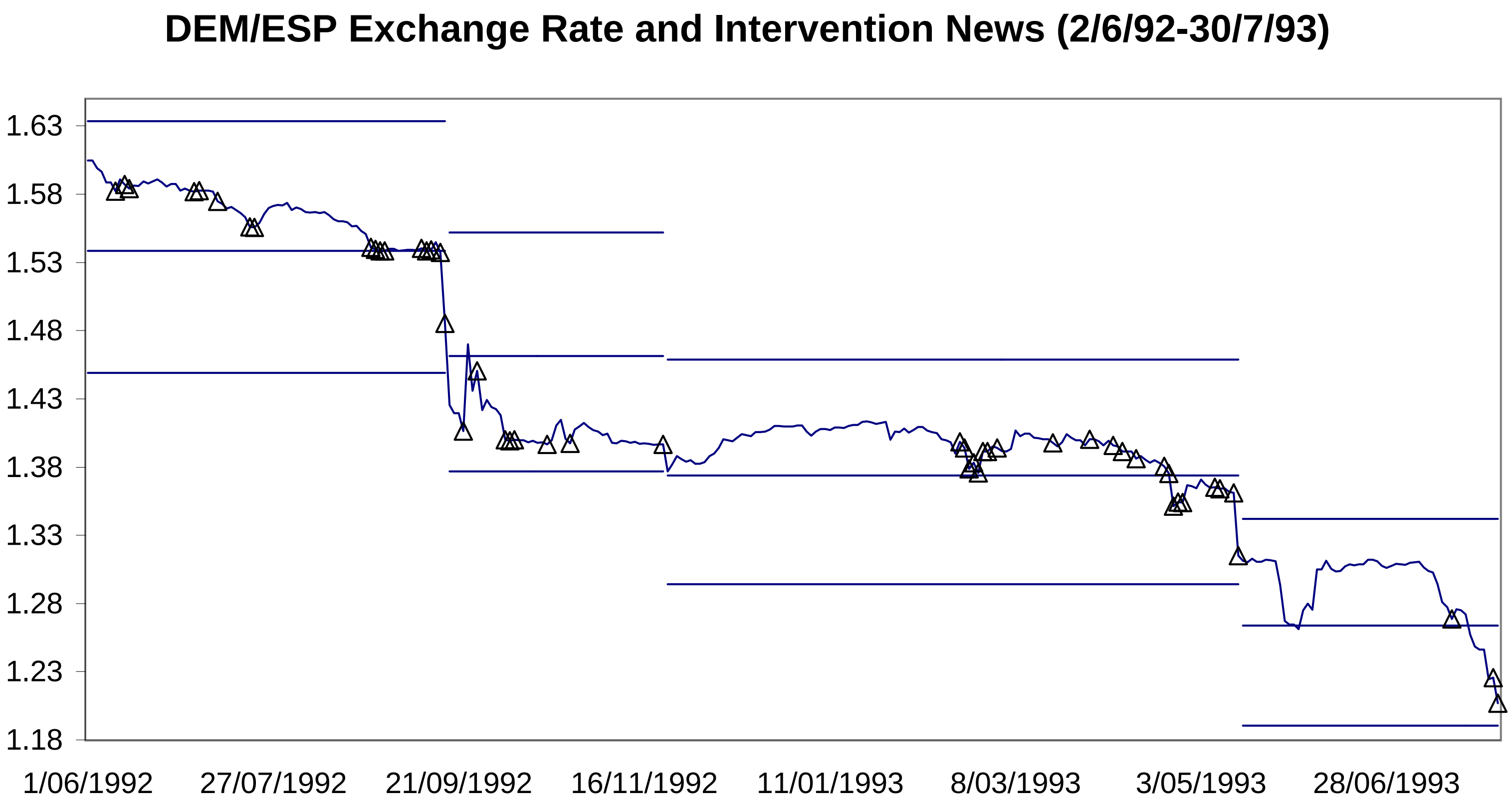
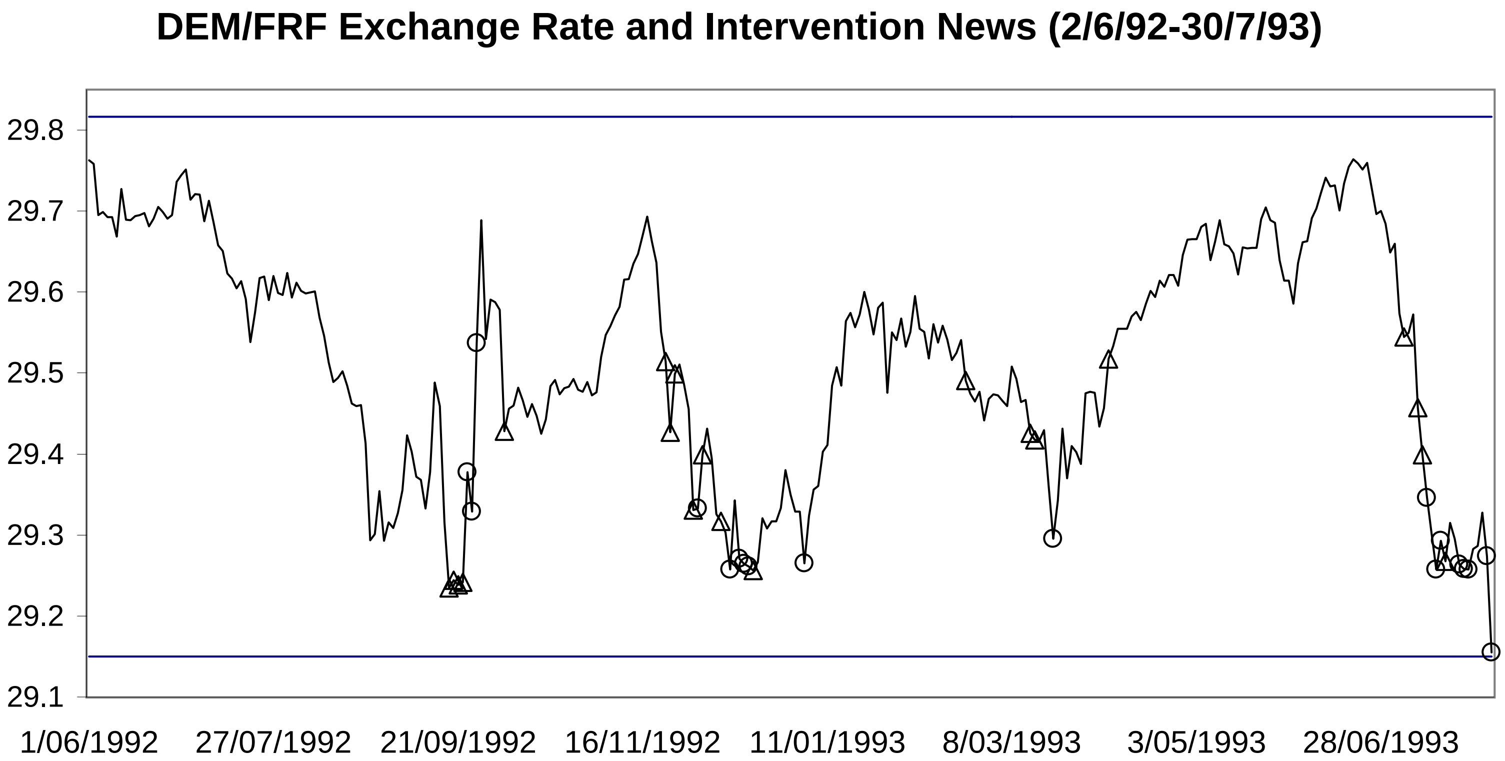


Figure 1: Daily Percentage Exchange Rate Changes, 2/6/92-30/7/93



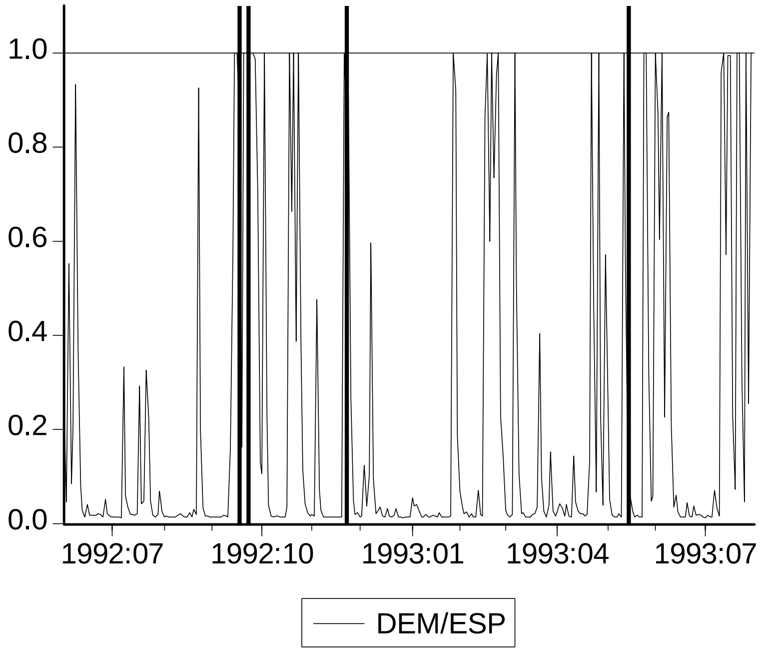
Note: The graph shows the DEM/ESP exchange rate, the DEM-central parity and its bilateral fluctuation band. The triangles indicate news about Banco de Espana intervention activities.

Figure 2: DEM/ESP Exchange Rate and Intervention News, 2/6/92-30/7/93



Note: The graph shows the DEM/FRF exchange rate, the DEM-central parity and its lower bilateral fluctuation band. The triangles indicate news about Banque de France intervention, the circles about joint Banque de France and Bundesbank intervention activities.

Figure 3: DEM/FRF Exchange Rate and Intervention News, 2/6/92-30/7/93



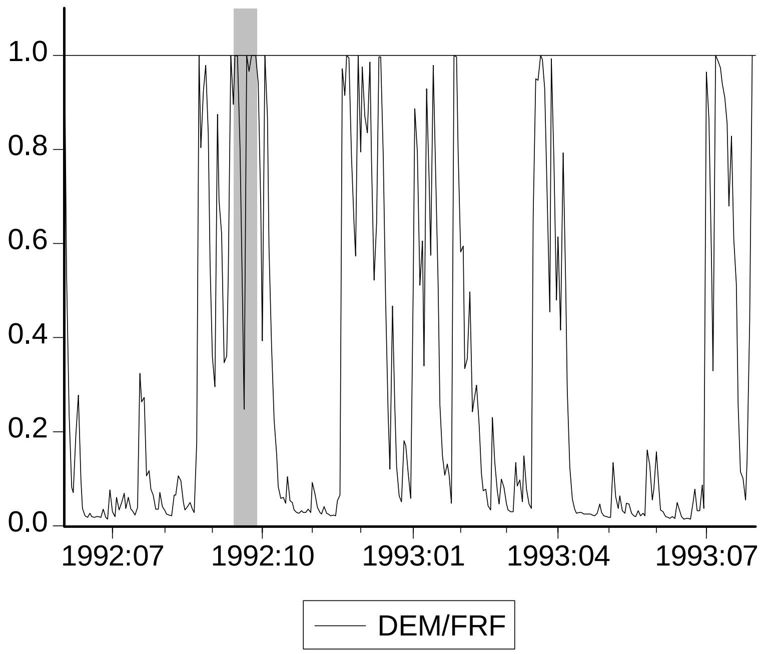


Figure 4: MS Model, Ex Post Probabilities of Regime 2, $P(Z_t = 2 \mid \Omega_t)$

Table 1: Descriptive Statistics for Daily Exchange Rates, 2/6/1992-30/7/1993

		ESP	FRF
Δs_t	Mean ($\times 10^2$)	-0.09	-0.01
		[0.01]	[0.40]
	Median ($\times 10^2$)	-0.03	0.00
	Maximum	4.44	0.71
	Minimum	-4.09	-0.51
	Std. Dev.	0.62	0.14
	Skewness	-1.00	0.07
	Kurtosis	23.89	7.16
	Jarque-Bera	5581.27	219.11
	$\rho(1)$	-0.01	-0.01
	$Q(20)$	21.99	16.64
		[0.34]	[0.68]
	$Q^2(20)$	89.13	118.66
		[0.00]	[0.00]]

Note: Δs_t is 100 times the log difference of the amount of Deutsche mark per 100 units of local currency (S_t). P-values in brackets. P-values below the means are based on a test of the null hypothesis of a zero mean. $\rho(1)$ denotes the value of the autocorrelation function at the first lag, $Q(20)$ and $Q^2(20)$ denote the Ljung-Box Q-statistics with lag length 20. The respective p-values are below the test statistics. The sample contains 304 daily observations.

Table 2: Intervention Frequency, 2/6/1992-30/7/1993

	ESP	FRF
Intervention News	50	38
Actual DEM-Sales	73	58
Public DEM-Sales	39	32
Secret DEM-Sales	34	26
BUBA Interventions	–	18

Source: Intervention News from Reuters News Service and Financial Times, actual intervention data from concertation protocols. "Intervention News" and "BUBA Interventions" refer to those interventions that were reported in Reuters News or in the Financial Times (irrespective of whether the news were true or false). "Actual DEM-Sales" are those interventions that were actually undertaken by the respective central banks. "Public DEM-Sales" are only those interventions take took place and that were publicly recognized.

Table 3: Markov Switching Estimation Results for ESP, 2/6/1992-30/7/1993

		Pref. Model (I)	Interv. News (II)	Sales (III)	Public Sales (IV)	Secret Sales (V)
Means	c_{10}	-2.88**	-3.10**	-2.93**	-2.98**	-2.76**
		(1.19)	(1.22)	(1.18)	(1.17)	(1.14)
	c_{20}	-33.04**	-33.18**	-34.17**	-34.24**	-32.60**
		(16.17)	(16.76)	(17.05)	(17.33)	(15.71)
Variances	σ_1^2	2.58**	2.65**	2.73**	2.67**	2.55**
		(0.51)	(0.48)	(0.58)	(0.56)	(0.46)
	σ_2^2	160.99**	166.03**	167.03**	168.77**	156.55**
		(55.32)	(58.00)	(57.71)	(59.80)	(50.65)
Transition	β_{10}	2.38**	2.71**	2.61**	2.49**	2.35**
Probabilities		(0.37)	(0.47)	(0.48)	(0.38)	(0.35)
	β_{11}		-1.62**	-0.80	-1.20**	3.19*
			(0.64)	(0.52)	(0.50)	(1.75)
	β_{20}	0.85**	0.70	0.38	0.38	0.78*
		(0.53)	(0.79)	(0.60)	(0.59)	(0.45)
	β_{21}		0.22	1.65**	1.50*	41.79
			(0.99)	(0.78)	(0.82)	(54.43)
Diagnostics	$\log L$	-79.64	-76.86	-75.63	-74.69	-77.89
	$LR-Stat$		5.57*	8.02**	9.90**	3.50
	$Q(20)$	14.01	15.88	14.46	18.32	12.18
	$Q^2(20)$	13.53	29.32*	26.92	11.39	12.51

Note: Standard errors (QML) in parentheses. A ** (*) means that the coefficient is significantly different from zero at a 5% (10%) level. N=304. The variables are defined on page 35. $Q(20)$ and $Q^2(20)$ denote the test statistics for the Ljung-Box Q-test with lag length 20. The null of no autocorrelation in the standardized residuals is rejected at a 5% (10%) significance level if the test statistic is greater than 31.41 (28.41). The null hypothesis for the likelihood ratio statistics is that β_{11} and β_{21} are jointly zero. The null is rejected at a 5% (10%) level if the test statistics is greater than 5.99 (4.61).

Table 4: Markov Switching Estimation Results for FRF, 2/6/1992-30/7/1993

		Pref. Model (I)	**Interv. News** (II)	**Sales** (III)	**Public Sales** (IV)	**Public S. w/o BUBA** (V)	**Secret Sales** (VI)
Means	c_{10}	0.17	0.03	-0.01	0.05	0.17	0.08
		-0.57	-0.64	-0.63	-0.61	-0.57	-0.58
	c_{20}	-24.05**	-24.15**	-25.74**	-25.77**	-25.38**	-23.57**
		-8.47	-9.41	-8.82	-8.8	-9.12	-8.9
	c_{21}	-36.43**	-37.31**	-39.6**	-39.56**	-38.9**	-35.86**
		-12.92	-14.19	-13.72	-13.8	-14.52	-13.44
Variances	σ_1^2	0.61**	0.61**	0.63**	0.62**	0.6**	0.61**
		-0.08	-0.08	-0.08	-0.09	-0.08	-0.07
	σ_2^2	4.35**	4.25**	4.39**	4.43**	4.5**	4.32**
		-0.94	-0.99	-0.98	-1.01	-0.99	-0.91
Transition	β_{10}	3.06**	3.41**	3.49**	3.26**	3.12**	3.53**
Probabilities		-0.42	-0.75	-0.58	-0.49	-0.42	-0.66
	β_{11}		-2.67	-1.08**	-0.77*	-0.54**	-5.35
			-1.99	-0.53	-0.43	-0.11	-4.14
	β_{20}	2.34**	2.67	2.16**	2.11**	2.29**	2.39**
		-0.75	-1.65	-0.82	-0.84	-0.78	-0.84
	β_{21}		-1.1	-0.02	-0.03	-0.23	2.26
			-2.24	-0.18	-0.17	-0.18	-3.38
Diagnostics	$\log L$	217.3	218.08	220.02	218.94	218.96	219.52
	$LR-Stat$		1.57	5.45*	3.28	3.32	4.44
	$Q(20)$	13.98	13.94	15.04	15.16	13.9	13.63
	$Q^2(20)$	21	14.77	14.05	15.52	15.26	19.35

Note: Standard errors (QML) in parentheses. A ** (*) means that the coefficient is significantly different from zero at a 5% (10%) level. N=304. The variables are defined on page 35. $Q(20)$ and $Q^2(20)$ denote the test statistics for the Ljung-Box Q-test with lag length 20. The null of no autocorrelation in the standardized residuals is rejected at a 5% (10%) significance level if the test statistic is greater than 31.41 (28.41). The null hypothesis for the likelihood ratio statistics is that β_{11} and β_{21} are jointly zero. The null is rejected at a 5% (10%) level if the test statistics is greater than 5.99 (4.61).

Table 5: Estimation Results for Expected Realignment Rate, 2/6/1992-30/7/1993

Dependent Variable: Expected Realignment Rate

$$g_t = c + \sum_{j=1}^{k} \gamma_j \cdot g_{t-j} + \alpha \cdot I_{t-1} + \beta \cdot (i_{t-1}^{3m} - i_{t-1}^{1m}) + \sqrt{h_t} \cdot \epsilon_t$$

ESP (OLS)

	DEM-sales	Public	Secret
I	1.01**		
	(0.47)		
II		1.65**	0.24
		(0.61)	(0.66)

FRF (EGARCH)

	Cond. Mean			**Cond. Variance**		
	DEM-sales	Public	Secret	DEM-sales	Public	Secret
I	0.28*			0.08**		
	(0.16)			(0.04)		
II		0.27**	0.41		0.08**	0.27
		(0.10)	(0.41)		(0.03)	(0.18)
	Public S. w/o BUBA		Secret	Public S. w/o BUBA		Secret
III		0.26	0.46		0.10**	0.13
		(0.14)	(0.47)		(0.04)	(0.22)

Note: The numbers represent the point estimates of the parameters (the $\hat{\alpha}$'s) of either DEM-sales or public and secret DEM-sales. "Public S. w/o BUBA" refers to those publicly recognized Banque de France intervention that were undertaken without Bndesbank support. For the DEM/ESP rate the estimates are based on an OLS regression and for the DEM/FRF rate on an EGARCH model. In addition to the parameters shown, the mean equation contains one lag (franc) and two lags of the dependent variable and realigment and capital control dummies for the peseta. Standard errors in parentheses. A ** (*) means that the coefficient is significantly different from zero at a 5% (10%) level. N=304. There is no evidence of remaining serial correlation in the residuals and squared standardized residuals.

Table 6: Estimated Reaction Functions, 2/6/1992-30/7/1993

$$I_t^* = c + \alpha \cdot D_{t-1} + \beta \cdot (h_{t-1} - h_{t-1}^*) + \gamma \cdot g_{t-1} + \epsilon_t \ ,$$

$$I_t = 0 \text{ if } I_t^* \leq 0 \qquad I_t = I_t^* \text{ if } I_t^* > 0$$

		α	β	γ
ESP	Sales	-1.93**	2.18**	0.03**
		(0.37)	(0.59)	(0.01)
	Public	-0.79	3.32**	0.05**
		(0.67)	(1.09)	(0.02)
	Secret	-2.27**	0.78	0.01
		(0.57)	(0.69)	(0.02)
	Public Regime 1	0.65	0.78	0.04**
		(0.47)	(0.85)	(0.01)
		$\hat{\alpha}$	$\hat{\beta}$	$\hat{\gamma}$
FRF	Sales	-20.32**	22.60	-0.32
		(7.80)	(25.86)	(0.23)
	Public	-36.54**	43.60	-0.39
		(13.32)	(38.61)	(0.39)
	Secret	1.67	-2.48	0.04
		(1.76)	(6.74)	(0.06)
	Public Regime 1	-6.43	29.98	0.05
		(6.44)	(26.80)	(0.23)
	Public S. w/o BUBA	-21.39	112.37**	-0.53
		(13.27)	(38.78)	(0.42)
	Public S. w/o BUBA Regime 1	-9.87	40.72	0.22
		(17.79)	(52.04)	(0.38)

Note: The numbers represent the point estimates of the Tobit reaction function shown above where I_t^* is the amount of either DEM-sales, public or secret DEM-sales. "Regime 1" refers to those intervention that took place in the first regime ("calm regime") of the Markov switching model. "w/o BUBA" refers to those publicly recognized Banque de France intervention that were undertaken without Bundesbank support. The equation for the DEM/ESP exchange rate also contains realignment and capital control dummies. Huber-White robust standard errors in parentheses. A ** (*) means that the coefficient is significantly different from zero at a 5% (10%) level. N=284. The variables are defined on page 35.

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