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Target zones and exchange rates: An empirical investigation

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Abstract

This paper develops an empirical model of exchange rates in a target zone. The distribution of exchange rate changes is conditioned on a latent jump variable where the probability and size of a jump vary over time as a function of financial and macroeconomic variables. When there is no jump, the target zone is credible and exchange rate changes are constrained to remain within the target zone band. The paper revisits the empirical evidence from the European Monetary System regarding the conditional distribution of exchange rate changes, the credibility of the system, and the size of the foreign exchange risk premia. In contrast to some previous findings, we conclude that the French Franc/Deutschmark rate exhibits considerable nonlinearities, realignments are somewhat predictable, and the credibility of the system did not increase substantially after 1987. Moreover, our model implies that the foreign exchange risk premium becomes large during speculative crises. © 1998 Elsevier Science B.V.

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1. Introduction

More than 20 years after the breakdown of the Bretton Woods system, the

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day-to-day behavior of exchange rates continues to puzzle both the academic community and policy makers. The recent currency crises in the European Monetary System (EMS) and the gyrations of the yen relative to the dollar, supposedly defying fundamentals, has heightened concerns among policy-makers about ''aberrant'' exchange rate movements. In fact, a recent IMF-study calls for the reimposition of managed target zones among the major currencies. These developments are likely to renew interest in the operation of exchange rate target zones. What are the dynamics of exchange rates and interest rates within target zones? To what extent have capital controls contributed to sustainable target zones and what triggers their breakdown? Are speculative crises of the sort recently encountered by the EMS predictable? Beginning with Krugman (1991), there is a large theoretical literature on target zones which has had limited empirical success when confronted with data from the EMS.

In this paper, we propose an empirical model of exchange rates that captures many salient features of target zones and allows some of the questions posed above to be addressed. In particular, our model of exchange rate changes accommodates occasional jumps of the type experienced in the EMS. Conditional on no jump occurring, we model exchange rate changes as being drawn from a truncated normal distribution with time-varying moments, so that there is zero probability of observing an exchange rate change which would take the exchange rate outside the target zone. That is, we model the system as being perfectly credible in the absence of jumps. Jumps can occur within the existing target zone or, in the event of a realignment, the exchange rate may move outside the current band. The mean size and variance of the jump, as well as its probability of occurring, vary over time as a function of financial and macroeconomic variables.

We focus on three empirical issues within our framework. First, our empirical model offers a rich characterization of the (conditional) distribution of the exchange rate.[1] Krugman's (Krugman (1991)) credible target zone model predicts that the exchange rate will exhibit (1) a nonlinear form of mean reversion and (2) conditional volatility that depends on the position of the exchange rate relative to the target zone boundaries. Whereas these attributes do not seem to be present in the data, our model enables us to filter out exchange rate jumps before examining these empirical predictions. Our volatility structure is more general than the credible-target-zone literature in two ways. First, in addition to dependence on the position of the exchange rate within the band, we incorporate generalized autoregressive conditional heteroscedasticity (GARCH) effects into our volatility model. Second, the total conditional volatility of exchange rate changes also incorporates the possibility of jumps. We refer to this component of volatility as *jump risk*.

[1]Previous studies focusing on the exchange rate distribution within a target zone include Beetsma, van der Ploeg, 1994, Chen, Giovannini, 1992, Engel, Hakkio, 1995, Flood et al., 1991 and Nieuwland et al., 1994.

Second, we provide an alternative measure of the credibility of a target zone. We define a target zone to be perfectly credible if the probability of the exchange rate moving outside the band is zero. Since we specify the complete conditional distribution of exchange rate changes, we are able to directly compute this probability, conditional on available information. The work of Svensson (1992a), (1992b), predicting small currency risk premia in target zones, has prompted many authors to rely on uncovered interest rate parity (UIRP) and interest differentials to infer market expectations of the exchange rate moving outside the band. We do not to take this route because evidence on UIRP is mixed (see Bossaerts, Hillion, 1991).

Third, we examine the currency risk premia implied by our model, providing an alternative test of the size of currency risk premia.

The remainder of the paper is organized as follows. The following section outlines the empirical model and its relationship to the current literature. Section 3 contains a discussion of estimation issues. The fourth section reports the empirical results and examines the conditional distribution of exchange rates. We apply the model to the French Franc/Deutschmark (FF/DM) rate from March 1979 until July 1993. During that period the FF and DM were part of the EMS. Section 5 focuses on the credibility of the target zone. The sixth section discusses implied currency risk premia and the final section concludes.

The main results can be summarized as follows. In contrast to previous empirical work such as Rose, Svensson (1995), we find evidence of nonlinearities in the FF/DM rate. We also find significant GARCH effects and time-varying jump risk in the conditional volatility process. Whereas the 1987–1991 period is often used as an example of a credible target zone (e.g., Ball, Roma, 1994), we find realignment probabilities to be quite substantial during that period. The fact that no realignments occur during a period does not imply that there is no realignment risk during that period. Our realignment probabilities do not increase dramatically ahead of the September currency crisis in 1992, but do spike up before the August 1993 crisis. Moreover, risk premiums tend to be large prior to realignments.

2. Motivation and econometric model

2.1. Theoretical target zone models

The theoretical target zone literature builds primarily on the stylized continuous-time model of Krugman (1991). The exchange rate, S_t, is measured in domestic currency per unit of foreign currency and is in logs. Moreover the exchange rate depends on market fundamentals and the expected exchange rate as implied by the simple monetary model of exchange rate determination. One fundamental is assumed to follow a Brownian motion and the other fundamental is controlled to keep the exchange rate within a pre-specified band. Two important assumptions underlie the model: the target zone is perfectly credible and it is defended with marginal interventions only.

The Krugman model has strong empirical implications for exchange rates and interest rate differentials which have been widely studied (see Svensson (1991a,b)). Of particular interest to us are the implications that the conditional volatility of exchange rate changes should be smaller near the edges of the band and the exchange rate should display a non-linear form of mean-reversion, even though fundamentals are Brownian motions.

Direct tests of the Krugman model in Smith, Spencer (1991) and De Jong (1994) have delivered clear rejections of the model. In particular, the endogenous nonlinearities are not sufficient to explain all of the leptokurtosis and ARCH effects in the data. Surprisingly, other studies such as Rose, Svensson (1995) and Flood et al., 1991 do not detect significant nonlinearities in EMS data. The history of repeated realignments and the preponderance of intra-marginal interventions in the EMS, described by Giavazzi, Giovannini (1989) and others, are also inconsistent with the Krugman model.

Two important extensions of the basic Krugman model are the introduction of realignment risk and intramarginal interventions. Bertola, Caballero (1992) introduce a (fixed) probability of a devaluation when the exchange rate hits the upper boundary. Bartolini, Bodnar (1992) show that this simple extension can generate realistic correlations between exchange rates and interest rate differentials. Ball, Roma (1993) and Bertola, Svensson (1993) generate realignments through an exogenous jump process with constant intensity.

To allow for intramarginal interventions, Delgado, Dumas (1991) and Lindberg, Söderlind (1993) introduce mean-reverting fundamentals and Beetsma, van der Ploeg (1994) stress the empirical importance of this extension. Lewis (1995) presents an alternative model with similar implications, where the mean reversion arises from occasional interventions by the authorities when the exchange rate deviates from target levels.

Our empirical model is inspired by all of these extensions to the Krugman model and is more general in several directions. This generality permits the derivation of a detailed set of stylized facts, which should prove useful for future theoretical work. In weighing the lack of theoretical underpinnings, we note that the monetary exchange rate model was virtually abandoned on empirical grounds (see Meese, Rogoff, 1983) before it was picked up by target zone theorists, mainly for convenience. In the remainder of this section, we discuss and motivate the most important features of our model.

2.2. An econometric target zone model

2.2.1. Incorporating Jumps

We begin by modeling exchange rates in a perfectly credible target zone, but also introduce the possibility of occasional jumps that may, or may not, take the exchange rate outside the band. One motivation for including the possibility of jumps is the work of Jorion (1988), who found evidence of jumps in floating exchange rates. Furthermore, in the EMS, discontinuities in the bilateral exchange rate may occur naturally for a number of reasons. First, and most obviously, a realignment of the EMS target zone may cause a jump in the exchange rate and is inevitable when the new and the old target zones do not overlap. Second, a jump in the exchange rate may occur within the target zone. For example, a pronounced change in the fundamental value of a currency can be caused by announcements of changes in central bank policy or by sudden speculative attacks on a weak currency. If the bilateral rate in question is in the lower half of the band when the announcement or attack occurs, there is room for a large and sudden depreciation in the bilateral rate to be accommodated within the band.

The reason we do not separately model within-the-band jumps and realignment jumps is threefold. First, both hedgers and speculators are concerned primarily with movements in exchange rates rather than movements in the EMS central parity. Our model focuses on the predictability of all jumps rather than just realignments, although we are able to separately examine the predictability of realignments. Second, there are relatively few realignment jumps, so including within-the-band jumps helps to identify the jump parameters. Finally, while the largest jumps in bilateral rates are associated with realignments, there are many within-the-band jumps which are of the same order of magnitude as the realignment jumps. Table 1 documents the ten largest decreases Table 2 and the

Table 1
The ten largest one-week decreases in the French Franc/Deutschmark (FF/DM) exchange rate (i.e., appreciation of the Franc) over the period 23 March 1979 to 23 July 1993, a total of 749 observations

Rank	Date	Percentage Change	Weeks Since Last Realignment[a]	Weeks Until Next Realignment[b]	Position in Band[c]
1	790615	−1.30	13	14	0.73
2	820205	−1.26	17	18	−0.36
3	920918	−1.22	296	–	0.86
4	800321	−1.17	25	80	−0.48
5	800516	−1.10	33	72	−0.05
6	810522	−1.08	86	19	0.89
7	861205	−0.91	34	5	0.72
8	810710	−0.80	93	12	0.71
9	831007	−0.57	28	130	0.04
10	810626	−0.56	91	14	0.62

[a]The number of weeks since the last realignment of the FF/DM target zone in the European Monetary System (EMS).
[b]The number of weeks between the current change and the next realignment of the FF/DM target zone in the EMS.
[c] The position of the FF/DM exchange rate within the EMS target zone before the current change. The difference between the current exchange rate and the center of the band as a proportion of half the width of the band. Values of −1, 0, and 1 correspond to exchange rates at the bottom edge, center, and top edge of the band, respectively.

Table 2
The ten largest one-week increases in the French Franc/Deutschmark (FF/DM) exchange rate (i.e., depreciation of the Franc) over the period 23 March 1979 to 23 July 1993, a total of 749 observations

Rank	Date	Percentage Change	Weeks Since Last Realignment[a]	Weeks Until Next Realignment[b]	Position in Band[c]
1	820611	5.61	35	0	0.97
2	811002	4.32	105	0	0.85
3	830318	3.52	39	0	0.91
4	860328	2.60	157	1	0.18
5	830304	2.05	37	2	0.05
6	820312	1.70	22	13	0.02
7	871023	1.68	40	–	−0.18
8	790608	1.45	12	15	0.09
9	810508	1.40	84	21	0.36
10	820212	1.29	18	17	0.91

[a]The number of weeks since the last realignment of the FF/DM target zone in the European Monetary System (EMS).
[b]The number of weeks between the current change and the next realignment of the FF/DM target zone in the EMS.
[c]The position of the FF/DM exchange rate within the EMS target zone before the current change. The difference between the current exchange rate and the center of the band as a proportion of half the width of the band. Values of −1, 0, and 1 correspond to exchange rates at the bottom edge, center, and top edge of the band, respectively.

ten largest increases in the FF/DM rate over the sample period. This period contains the six realignments of the FF/DM central parity that have occurred since the inception of the EMS. While three of those realignments drive the three largest changes in the bilateral rate, the other three realignments do not rank in the top ten changes in the bilateral rate. That is, several within-the-band jumps are larger than several realignment jumps. Two of the within-the-band jumps appear to be associated with speculative attacks leading up to a realignment. Over the sample period, jumps in the bilateral rate are more likely to take the form of a depreciation of the Franc than an appreciation of the Franc. There are nine increases in the FF/DM rate that are larger than the largest decrease in the FF/DM rate.

Formally, we seek to model the distribution of log changes in exchange rates, conditional on available information, $f(\Delta S_t | I_{t-1})$, where I_{t-1} denotes an information set, and $f(\cdot | \cdot)$ denotes a conditional density. We construct this conditional distribution by separately modeling (1) the conditional distribution in the absence of jumps, and (2) jumps in the exchange rate. To this end, we define an indicator variable:

$$J_t = \begin{cases} 1 & \text{if the exchange rate jumps at time } t \\ 0 & \text{otherwise} \end{cases}$$

The conditional distribution of exchange rate changes can be factored as:

$$f(\Delta S_t|I_{t-1}) = f(\Delta S_t|I_{t-1},J_t = 0)Pr(J_t = 0|I_{t-1}) + f(\Delta S_t|I_{t-1},J_t = 1)Pr(J_t = 1|I_{t-1}) \quad (1)$$

Note that in contrast to Ball, Roma (1993), Bertola, Caballero (1992), and Bertola, Svensson (1993), we allow for *jumps*, rather than just *realignments*. Engel, Hakkio (1995) use a model that is similar to ours, however they model J_t as a Markov process and do not take the presence of the band explicitly into account (see also Avesani, Gallo, 1996). Below, we describe the precise specification of the component pieces of our density and we summarize the full model in Table 3. The precise definitions of the variables used in the model are summarized in Table 4.

2.2.2. A Model of a Credible Target Zone

First, consider $f(\Delta S_t|I_{t-1},\ J_t = 0)$, the distribution of exchange rate changes conditional on available information, and on there being no jump. This corresponds to the setting of a properly functioning target zone system where the exchange rate will never move outside the target zone, in which case exchange rate changes are bounded. The maximum possible change, Δ_U, takes the exchange rate to the upper boundary of the target zone, and the minimum possible change, Δ_L, takes the exchange rate to the lower boundary. In such a system, a bounded density is required to model exchange rate changes. In this paper, we use a truncated normal density which has a very flexible form, few parameters to estimate, and is defined only on $[\Delta_L,\ \Delta_U]$.

The truncated normal distribution is:

$$f(\Delta S_t|I_{t-1},\ J_t = 0) = \frac{\phi\left(\dfrac{\Delta S_t - \mu_{t-1}}{\sqrt{\sigma^2_{t-1}}}\right)\dfrac{1}{\sqrt{\sigma^2_{t-1}}}}{\Phi\left(\dfrac{\Delta_{U_{t-1}} - \mu_{t-1}}{\sqrt{\sigma^2_{t-1}}}\right) - \Phi\left(\dfrac{\Delta_{L_{t-1}} - \mu_{t-1}}{\sqrt{\sigma^2_{t-1}}}\right)}$$

Table 3
Model Specification

The complete model is

$$f(\Delta S_t|I_{t-1}) = \begin{cases} TN(\mu_{t-1},\sigma^2_{t-1},\Delta_{L_{t-1}},\Delta_{U_{t-1}} & \text{w.p. } (1-\lambda_{t-1}) \\ N(\rho_{t-1},\rho^2_{t-1}\delta^2) & \text{w.p. } \lambda_{t-1} \end{cases} \quad (10)$$

where

$\lambda_{t-1} = \Phi(\beta_1 + \beta_2 SYC_{t-1})$
$\rho_{t-1} = \beta_3 + \beta_4 LR_{t-1} + \beta_5|PB_{t-1}| + \beta_6 ID_{t-1} + \beta_7 CID_{t-1}$
$\mu_{t-1} = \beta_8 + \beta_9 PB_{t-1}$
$\sigma^2_{t-1} = \beta_{10} + \beta_{11}(1 - RD_{t-1})\epsilon^2_{t-1} + \beta_{12}\sigma^2_{t-1} + \beta_{13}|PB_{t-1}|,$

The precise definitions of the variables used in the model are summarized in Table 4. *TN* indicates a truncated normal distribution where the underlying normal distribution has conditional mean μ_{t-1} and conditional variance σ^2_{t-1}. λ_{t-1} is the probability of a jump ($J_t = 1$) and $\Phi(\cdot)$ denotes the cumulative normal distribution function. RD_t is a realignment dummy variable which takes the value one when a realignment of the target zone occurred in week t and zero otherwise.

Table 4
Definitions of the Variables Used in the Model

Exchange Rate Changes: ΔS_t
Continuously compounded exchange rate change: $\ln\left(\frac{S_t}{S_{t-1}}\right)$ where S_t represents the French Franc/Deutschmark exchange rate at time t.
Position In Band: PB_t
The relative position of the French Franc/Deutschmark (FF/DM) exchange rate within the European Monetary System (EMS) target zone band: $\frac{S_t - C_t}{\frac{1}{2}(U_t - L_t)}$. S_t is the FF/DM exchange rate, C_t is the center of the EMS target zone, U_t is the upper boundary of the target zone, and L_t is the lower boundary of the target zone. $-1<PB_t<1$, with $PB_t>0$ if the franc is in the weak half of the band.
Level of Reserves: LR_t
The level of foreign currency reserves of the Banque de France, relative to a four-week moving average: $\frac{\text{Reserves}_t}{\frac{1}{4}\Sigma_{i=1}^{4} Reserves_{t-i}}$. When $LR_t<1$, foreign currency reserves have been depleted relative to their recent average level.
Slope of Yield Curve: SYC_t
The slope of the yield curve for French franc-denominated instruments: $i_t^{\mathrm{F12}} - i_t^{\mathrm{F1}}$, where i_t^{F12} is the one-year rate and i_t^{F1} is the one-month rate and both rates are nominal eurocurrency yields.
Cumulative Inflation Differential: CID_t
The cumulative inflation differential between France and Germany: $\frac{CPI_t^{\mathrm{F}}}{CPI_{0t}^{\mathrm{F}}} - \frac{CPI_t^{\mathrm{G}}}{CPI_{0t}^{\mathrm{G}}}$. CPI_{0t}^{F} represents the CPI level in France at the time of the last realignment of the French Franc/Deutschmark target zone and CPI_t^{F} represents the *CPI* level in France at time *t*. The corresponding terms relate to the *CPI* level in Germany. This variable measures the difference between inflation in France and inflation in Germany since the most recent realignment.
Interest Differential: ID_t
The interest differential between France and Germany: $i_t^{\mathrm{FW}} - i_t^{\mathrm{GW}}$. This is the difference between one-week Eurocurrency rates in France and Germany.

where $\phi(\cdot)$ denotes the standard normal probability density function and $\Phi(\cdot)$ represents the standard normal cumulative distribution function and μ_{t-1} and σ_{t-1}^2 are the conditional mean and variance of the normal distribution which is to be truncated. That is, ΔS_t is modeled as being normally distributed with conditional mean μ_{t-1} and variance σ_{t-1}^2, with any probability mass falling outside the range of $[\Delta_L, \Delta_U]$ being truncated and added back in proportion to the density within this range.

We parameterize the conditional mean, μ_{t-1}, to incorporate mean reversion by letting exchange rate changes depend on the position in the band: $\mu_{t-1} = \beta_8 + \beta_9 PB_{t-1}$, where PB_{t-1} takes a value on $[-1, 1]$ indicating the relative position of the exchange rate within the target zone. When $PB_t = 0$, the exchange rate is at the center of the target zone, when $PB_t = 1$, the exchange rate is at the upper boundary of the target zone, and so on. Hence, the expected change toward the center of the target zone is stronger when the exchange rate is near the edge of the band. This is consistent with the predictions of a Krugman-type model and also with the presence of intra-marginal interventions. For example, central banks may defend an implicit exchange rate target within the band. Although there is no official record of such an implicit target for the French Franc, narrower bands were maintained by both the Dutch and Belgian central banks for parts of the sample period.

The conditional variance, σ^2_{t-1}, follows a GARCH (1, 1) process,[2] augmented to allow for dependence on the position within the band (which is the sole determinant of the conditional variance in a standard target zone model):

$$\sigma^2_t = \beta_{10} + \beta_{11}(1 - RD_{t-1})\epsilon^2_{t-1} + \beta_{12}\sigma^2_{t-1} + \beta_{13}|PB_{t-1}| \tag{2}$$

In this specification, $\epsilon_t = \Delta S_t - E_{t-1}[\Delta S_t]$, as is standard, and RD_t is a realignment dummy variable which takes the value 1 when a realignment of the target zone occurred in week t and zero otherwise. This variable is included to capture the effects of ''pressure-relieving'' shocks. It is common in the EMS for realignments to be preceded by periods of above-average volatility, often caused by speculative attacks and fears of a sudden depreciation in the value of a currency. The period immediately after a realignment is usually characterized by below-average volatility, as the weak currency's competitive position has been restored, the exchange rate is usually near the center of the band, and the probability of another realignment in the near future is small. Since realignments often cause a large one-time shock to the exchange rate, ϵ^2_{t-1} is large and a standard GARCH model would predict very high volatility after a realignment—the opposite of what is expected. We therefore model realignment shocks as being non-persistent.

2.2.3. Modeling Jumps

Next, we consider the possibility that the exchange rate may jump, in which case the relevant piece of the density is $f(\Delta S_t|I_{t-1}, J_t = 1)$. Since there is no a priori reason to impose an upper or lower limit on the magnitude of a jump, we model jumps in the exchange rate as being drawn from a normal distribution. We allow the moments of the jump distribution to be state-dependent (i.e. dependent on I_{t-1}, the information set). The form of this state-dependence renders the conditional mean proportional to the conditional standard deviation. For example, if the conditional mean of the jump distribution is small and positive we constrain the conditional variance to be small. This assumption limits parameter proliferation and avoids identification problems in situations where a positive jump is expected (the conditional mean jump size is positive), while at the same time there is a significant probability of a negative jump (the conditional variance of the jump size is large). In particular, when the exchange rate jumps, we model changes in exchange rates (ΔS_t) as being (conditionally) normally distributed with conditional

[2] See Bollerslev et al. (1992) for a review of evidence on the presence of GARCH effects in a variety of asset prices.

mean ρ_{t-1} and conditional variance $\rho_{t-1}^2\delta^2$ where δ is a scaling parameter to be estimated and

$$\rho_{t-1} = \beta_3 + \beta_4 LR_{t-1} + \beta_5|PB_{t-1}| + \beta_6 ID_{t-1} + \beta_7 CID_{t-1}$$

These conditioning variables, which are defined precisely in Table 4, are:

1. The difference between the level of French reserves and a moving average of previous reserve levels, LR_{t-1}. Although in the EMS intervention is only mandatory when the exchange rate hits the boundary of the band, central banks that foresee speculative pressures, might, and do, intervene intramarginally.[3] Therefore, changes in the level of reserves might signal a higher probability of either a realignment or a swift movement to the edge of the band. Speculative attacks may also drive reserves to a critically low level beyond which the target zone becomes unsustainable and a realignment becomes inevitable.[4] We use the moving average specification for two reasons. First, the level of reserves is measured with error. By measuring reserves relative to a moving average, we capture the depletion of reserves that occurs during a speculative crisis without regard to the level of reserves. Second, it is common for reserves to be depleted quite dramatically in the month before the realignment (see Collins, 1992). Therefore, examining reserves relative to a moving average is likely to provide a strong signal of impending realignments.
2. The position of the exchange rate within the band, $|PB_{t-1}|$. Larger jumps are expected near the edges of the band. At the lower boundary, a larger jump can be accommodated within the band, and at the upper boundary, the only kind of jump that is possible is a realignment jump, which tend to be relatively large.
3. The interest differential with German interest rates, ID_{t-1}. Speculative tensions and the ensuing actions of monetary authorities to defend the currency are reflected rapidly in interest rates (see Svensson, 1991b). In particular, the yield curve typically inverts and the differential with the German interest rate increases when a devaluation of the French Franc is expected. Both the slope of the yield curve and short-term interest rate differentials can serve as jump predictors, and they are highly correlated. In this model, we try to disentangle the size and probability of jumps which requires that size and probability do not depend on the same set of instruments. When UIRP holds, the interest rate differential equals expected exchange rate changes reflecting both the expected

[3]Giavazzi, Giovannini (1989) discuss the evidence of intra-marginal interventions within the EMS. Note, however, that the Bundesbank has never intervened intra-marginally.

[4]Bertola, Caballero (1991) present a stylized Krugman type model in which the realignment probability depends on the level of reserves. The relationship between the large literature on speculative attacks on fixed rate systems and target zone models is explored in Flood, Garber (1991). In Eichengreen et al. (1995) the level of reserves is an important part of an ‘‘index of speculative pressure’’.

size and probability of jumps. Empirically, however, the slope of the yield curve is a better jump predictor while the interest differential is a better jump size indicator (see below).

4. The cumulative inflation differential between Germany and France since the most recent realignment, CID_{t-1}. Larger jumps are expected when CID is large. Since Germany has consistently achieved lower inflation than most other EMS countries, an unchanged EMS-band gradually erodes the competitiveness of these countries as their real exchange rates appreciate. In the early years of the EMS, the realignments typically restored competitiveness. From 1983 onwards, however, realignments, although still highly correlated with cumulative inflation differentials, no longer fully compensated for lost competitiveness so that inflationary policies were punished (see Giavazzi, Pagano, 1988). Our model allows us to assess whether the cumulative inflation differential has any ex-ante bearing on the size of jumps.

2.2.4. The conditional probability of Jumps

The only piece of the model left to be specified is the time-varying probability of a jump, $\lambda_{t-1}=Pr(J_t=1|I_{t-1})$. Since this is a true probability, we constrain $0<\lambda_{t-1}<1$ using the normal cumulative distribution function, as in a probit model. We let the jump probability be a function of the slope of the yield curve, SYC_{t-1}:$\lambda_{t-1}=\Phi(\beta_1+\beta_2 SYC_{t-1})$. It is likely that the jump probability is influenced by a number of macroeconomic variables. Within the band, monetary policy has some independence of pursuing other goals. When macroeconomic developments, such as poor GNP-growth and high unemployment, create tensions between exchange rate and other macro-economic goals, the jump probability may rise. However, the relationship between macro-economic data and jump probabilities is likely to be noisy and it is difficult to construct weekly macro-economic data that were actually in the information sets of economic agents. The slope of the yield curve, through the forward looking nature of market-determined interest rates, may better reflect such effects than poorly measured macroeconomic data.

Of course, the variables that determine the mean and variance of the jump, described above, may also affect the jump probability. However, we found them to be statistically and economically insignificant jump predictors in the presence of the slope of the yield curve variable.[5]

3. Estimation

The model described above is written in terms of the conditional distribution of exchange rate changes and is a reduced-form model. Full information maximum likelihood requires that we model the joint density of exchange rate changes *and*

[5]For example, when λ_{t-1} is allowed to depend on the level of reserves, its coefficient is insignificantly different from zero and of the wrong sign.

the conditioning variables, which is beyond the scope of this paper. In empirical work in this area, it is customary to proceed by maximizing the *conditional* likelihood function $\Sigma_{t=2}^{T} L(\Delta S_t | I_{t-1})$. While this approach produces consistent estimates of the parameters of our model, some degree of efficiency is sacrificed by parameterizing only part of the full likelihood function. To see this, define Z_t to be a vector consisting of all of our conditioning variables, and let $\tilde{Z}_t' = \{Z_t', Z_{t-1}', ..., Z_1'\}$. Let $\Delta\tilde{S}_t$ be defined in similar fashion. The information set is made up of these two components so that $I_t = \{\Delta\tilde{S}_t, \tilde{Z}_t\}$. Finally, define θ to be the vector of parameters affecting the joint distribution $f(\Delta\tilde{S}_t, \tilde{Z}_t)$.

Full maximum likelihood estimation requires maximization of the likelihood function based on the joint density of the data and the conditioning variables:

$$L(\Delta\tilde{S}_T, \tilde{Z}_T; \theta) = \ln[f(\Delta\tilde{S}_T, \tilde{Z}_T; \theta)] \tag{3}$$

A series of conditioning arguments can be used to establish that, up to an initial condition,

$$f(\Delta\tilde{S}_T, \tilde{Z}_T; \theta) = \Pi_{t=1}^{T} f(Z_t | \Delta S_t, I_{t-1}; \theta) f(\Delta S_t | I_{t-1}; \theta) \tag{4}$$

Conditioning also on our jump variable, J_t, yields

$$f(\Delta\tilde{S}_T, \tilde{Z}_T; \theta) = \Pi_{t=1}^{T} \sum_{i=0}^{1} f(Z_t | S_t, J_t = i, I_{t-1}; \theta) f(\Delta S_t | J_t = i, I_{t-1}; \theta) \\ Pr(J_t = i | I_{t-1}) \tag{5}$$

We then assume that once we condition on the exchange rate, the conditioning variables are independent of the jump variable. That is

$$f(Z_t | \Delta S_t, J_t = i, I_{t-1}; \theta) = f(Z_t | S_t, I_{t-1}; \theta) \tag{6}$$

Clearly, the conditioning variables are not independent of the contemporaneous exchange rate changes in our setting. For example, when ΔS_t is large as a result of a speculative attack, it is quite likely that LR_t and SYC_t will be low and ID_t will be high. The assumption that is made above merely posits that after conditioning on the change in exchange rates, the jump variable is not informative about the distribution of the conditioning variables. That is, knowledge of the actual change in exchange rates is at least as informative as knowing whether there was a jump in exchange rates. If a large change in the exchange rate is observed, it is not possible to determine (ex post) whether this was the result of a jump or a draw from the tail of the no-jump distribution. Therefore, the effect of this change in the exchange rate in the model is the same, regardless of the cause. Alternatively, a realignment is observable (ex post) since the change in the target zone boundaries will appear in the information set, and the dynamics of the variables change immediately after a realignment. Even in this case, however, conditioning on the jump variable (J_t) is uninformative when the change in exchange rate (ΔS_t) and target zone boundaries ($\Delta_{L_{t-1}}$) and ($\Delta_{U_{t-1}}$) are observable. This is because when

ΔS_t moves the exchange rate outside the existing band, it can only be due to a realignment.[6] This allows us to write the log-likelihood function as

$$L(\Delta S_T, \tilde{Z}_T; \theta) = \sum_{t=2}^{T} \ln[f(Z_t|\Delta S_t, I_{t-1}; \theta_2)] + \sum_{t=2}^{T} \ln\{\sum_{i=0}^{1} f(\Delta S_t|J_t = i, I_{t-1}; \theta_1) Pr(J_t = i|I_{t-1})\}, \quad (7)$$

where θ_1 is the vector of parameters affecting the conditional distribution of exchange rate changes and θ_2 is the vector of parameters affecting the conditional distribution of the instruments. We proceed by parameterizing only the second piece of this log-likelihood function. Since this second piece allows identification of θ_1, our maximum likelihood estimates will be consistent. While the first piece of the log-likelihood function may contain potentially relevant information, parameterization of this joint density is not feasible given the number of conditioning variables and the size of our data set. Since consistent estimates of all of our parameters can be obtained by focusing exclusively on the second piece, this is the procedure we employ.

4. Results

4.1. Data

Our data were obtained from *Datastream, The Financial Times*, and the International Monetary Fund publications *Exchange Arrangements* and *International Financial Statistics*. The sample consists of weekly data from March 1973, with the start of the Exchange Rate Mechanism (ERM) within the EMS, through to July 1993, a total of 749 observations. We use Euro-currency interest rates in the empirical analysis because they are true market-determined rates. The presence of capital controls in France, for most of the sample period, implies that there can be a significant wedge between domestic rates and true market rates. Fig. 1 plots the FF/DM exchange rate and EMS bounds over the sample period.

4.2. Results

The econometric model in Table 3 was estimated by maximum likelihood using the GAUSS MAXLIK and CML modules. The parameter estimates reported have been verified by using two optimization algorithms (Berndt, Hall, Hall, and

[6]Note that this factorization would not be valid if J_t followed a persistent Markov process as in Engel, Hakkio (1995). Gray (1996b) discusses the problems in constructing the likelihood function for regime-switching models with state-dependent transition probabilities.

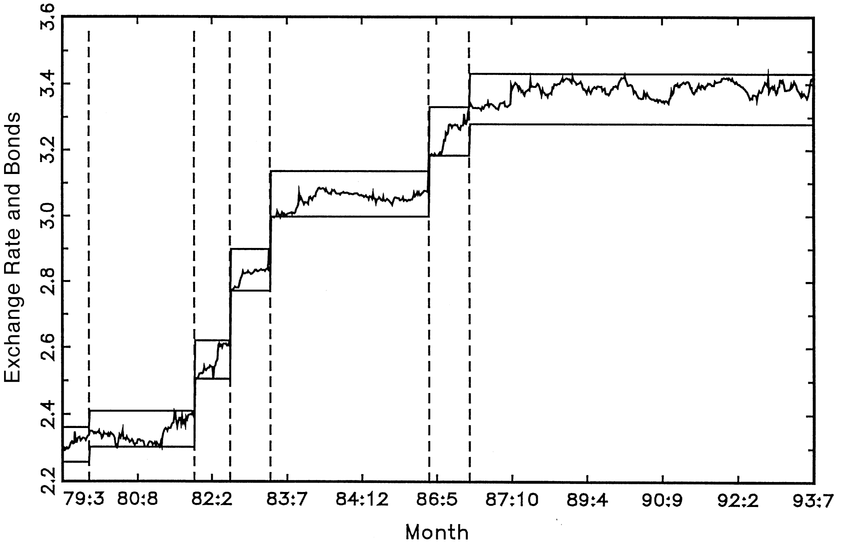

Fig. 1. French Franc/Deutschmark (FF/DM) spot rates and European Monetary System target zone boundaries. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

Hausman and Broyden, Fletcher, Goldfarb, and Shanno), and several different sets of starting values. While the likelihood functions are nonlinear and may admit several local optima, this procedure provides some degree of confidence that the parameter estimates reported below are at the global maximum. In all cases, asymptotic standard errors are heteroskedasticity-consistent. We also report results for a model in which we use normal distributions for both $J_t = 1$ (jumps) and $J_t = 0$ (no jumps) (Model 2). We only use this model to provide for a comparison with standard models, which ignore the possibility of jumps and are based on a single conditionally normal distribution.

The parameter estimates are reported in Table 5. The model confirms much of the intuition reviewed above, with most parameters taking the hypothesized sign and six reaching statistical significance. In Table 6, we report a number of likelihood ratio tests (LRT) of restricted models which are discussed below.

4.2.1. Mean Reversion Within the Band

In our truncated normal model, the mean-reversion parameter β_9 is negative, but not statistically significant according to the t-test or the LRT. In the truncated normal distribution, however, $\mu_{t-1} = \beta_8 + \beta_9 \; PB_{t-1}$ represents the conditional mean of the underlying normal distribution which is truncated, and not the

Table 5
Maximum Likelihood Parameter Estimates

Parameter	Model 1	Model 2
β_1	−1.2891*	−1.2046*
	(0.1607)	(0.1666)
β_2	−0.0129	−0.0150
	(0.0120)	(0.0133)
β_3	0.8111*	0.7414*
	(0.3204)	(0.2974)
β_4	−0.6264*	−0.5719*
	(0.2772)	(0.2584)
β_5	−0.0896	−0.0740
	(0.0868)	(0.0800)
β_6	0.0371*	0.0329*
	(0.0126)	(0.0117)
β_7	0.8924	0.8665
	(0.7830)	(0.7440)
β_8	−0.0099	−0.0049
	(0.0095)	(0.0100)
β_9	−0.0157	−0.0476*
	(0.0195)	(0.0169)
β_{10}	0.0031	0.0040
	(0.0022)	(0.0030)
β_{11}	0.2957*	0.2523*
	(0.1060)	(0.0936)
β_{12}	0.3514*	0.3904*
	(0.1639)	(0.1723)
β_{13}	0.0190	0.0121
	(0.0102)	(0.0081)
δ^2	2.8104	3.2715
	(1.5836)	(1.7731)

The sample contains weekly data from 23 March 1979 to 23 July 1993, a total of 749 observations. Model 1 denotes the model in which exchange rate changes are assumed to be conditionally distributed as a truncated normal, with conditionally normal jumps. Model 2 denotes the model in which exchange rate changes are assumed to be conditionally normal, with conditionally normal jumps. * = significant at the 5% level. Model 1 is described in Table 3.

conditional mean of the resulting truncated distribution. When the exchange rate is near the top of the band, the right half of the distribution is truncated more severely than the left, leaving a negatively skewed distribution. Whereas μ_{t-1}, defined above, denotes the peak of this distribution, the mean of the distribution will be lower, further towards the center of the band. Symmetrically, when the exchange rate is near the lower boundary, the left half of the distribution will be more severely truncated, resulting in positive skewness. More formally, the conditional mean of the truncated distribution is:

Table 6
Likelihood ratio tests of restricted models

Test Number	Description of Test	Restriction	Degrees of Freedom	LRT Statistic [p-Value]
1	Mean Reversion	$\beta_9=0$	1	0.7145 [0.3979]
2	GARCH Volatility	$\beta_{11}=\beta_{12}=0$	2	65.0721 [0.0000]
3	Position in Band Volatility	$\beta_{13}=0$	1	11.4128 [0.0007]
4	Conditional Heteroskedasticity	$\beta_{11}=\beta_{12}=\beta_{13}=0$	3	71.8435 [0.0000]
5	Time-Varying Jump Size	$\beta_4=\beta_5=\beta_6=\beta_7=0$	4	39.6978 [0.0000]
6	Time-Varying Jump Probability	$\beta_2=0$	1	1.5789 [0.2089]
7	Time-Varying Jump Size and Probability	$\beta_2=\beta_4=\beta_5=\beta_6=\beta_7=0$	5	48.6742 [0.0000]
8	Existence of Jumps	$\beta_1=\ldots=\beta_7=0$	7	646.4021 [0.0000]

Test 1 examines the significance of the dependence of the conditional mean on the position in the band in the absence of jumps. Recall, however, from Fig. 2, that the shape of the truncated normal distribution itself can drive mean reversion within the bands. Test 2 examines the significance of GARCH effects in the conditional volatility. Test 3 examines whether the conditional volatility depends on the position in the band. Test 4 examines the joint significance of both sources of conditional heteroskedasticity. Test 5 examines the joint significance of the conditioning variables in predicting the jump size. Test 6 examines the significance of allowing for time variation in the jump probability. Test 7 jointly examines time variation in the jump size and probability. Test 8 examines the existence of jumps. For this final test, the LRT statistic is not necessarily distributed as a χ^2 under the null, since the jump size parameters are not identified under the null of no jumps.

$$E[\Delta S_t | I_{t-1}, J_t = 0] = \mu_{t-1} + \frac{\phi\left(\dfrac{\Delta_{L_{t-1}} - \mu_{t-1}}{\sigma_{t-1}}\right) - \phi\left(\dfrac{\Delta_{U_{t-1}} - \mu_{t-1}}{\sigma_{t-1}}\right)}{\Phi\left(\dfrac{\Delta_{U_{t-1}} - \mu_{t-1}}{\sigma_{t-1}}\right) - \Phi\left(\dfrac{\Delta_{L_{t-1}} - \mu_{t-1}}{\sigma_{t-1}}\right)} \sigma_{t-1} \tag{8}$$

Fig. 2 plots this conditional expectation for each observation in the sample, ordered by the position of the exchange rate within the target zone, at the unconditional variance. The mean reversion is evident from the fact that when the exchange rate is close to the boundaries, movements of over 0.5% are expected. An expected one-week change in exchange rates of this magnitude is clearly of economic significance. To compute confidence bands, we generate 1000 samples of parameter values from the multivariate asymptotic distribution of parameters. For each data point, we recompute the expected change in exchange rates (the

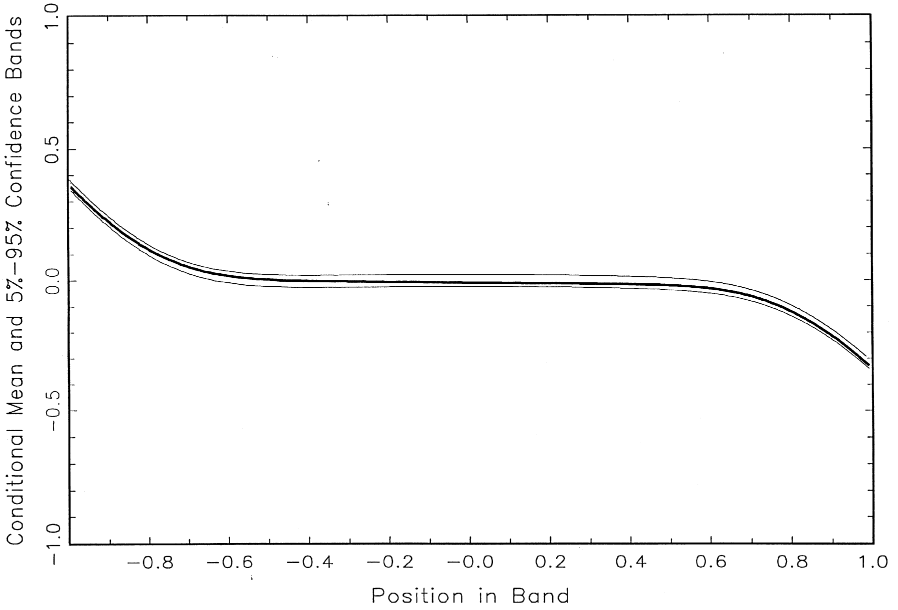

Fig. 2. This figure shows the expected change in the French Franc/Deutschmark exchange rate over the following week—conditional on available information and on no jump occurring—as a function of the position of the exchange rate within the European Monetary System target zone. This is a measure of reversion towards the center of the band. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

mean reversion) for each set of parameters, yielding 1000 estimates of the expected change in exchange rates. The confidence bands are then taken as the 5th and 95th percentile of the 1000 estimates. Clearly the confidence bands are quite tight, indicating that the mean reversion is also of statistical significance.

These results are consistent with intramarginal central bank intervention, and are inconsistent with the regulated Brownian motion assumption of the standard Krugman model where interventions occur only at the edges of the band.[7] Nevertheless, the non-linear nature of the mean reversion is broadly consistent with the predictions of Krugman-type models. Svensson (1992b) reports evidence inconsistent with this prediction of many theoretical target zone models by examining the relationship between the interest differential and the position of the exchange rate within the band. This evidence, however, depends on two auxiliary assumptions not imposed on our model: uncovered interest rate parity and a fully credible target zone. Note also, that in Model 2 where within-the-band changes are

[7]They may also reflect the existence of an effective narrower bound in a multilateral target zone system as in Pill (1994) or the presence of externality effects in such a multilateral system as in Flandreau (1996).

modeled as being normally distributed, β_9 is significantly less than zero, indicating strong mean reversion.

4.2.2. Conditional Heteroskedasticity

In contrast to the predictions of Krugman-type models, we find that the conditional volatility of exchange rates does not decrease as the exchange rate approaches the boundaries of the target zone. Our maximum likelihood estimates of β_{13} are positive and although the *t*-statistic does not reach significance, the more powerful LRT does indicate a significant *increase* in volatility as the band is approached. Moreover, there are also highly significant GARCH effects which is also inconsistent with the assumption of regulated Brownian motion in Krugman-type models. β_{11} and β_{12} are individually significant according to the *t*-tests in Table 5 and jointly significant according to the LRT in Table 6. Not surprisingly, the LRT for the joint significance of both sources of conditional heteroscedasticity indicates strong significance.

Furthermore, the persistence of conditional variance shocks has been reduced by the introduction of jumps and dependence on the position of the exchange rate in the band, and by modeling realignment shocks as being non-persistent. In our model, the effects of conditional variance shocks die out relatively quickly, with $\beta_{11}+\beta_{12}=0.6471$. Contrast this with (1) a standard GARCH (1, 1) model where $\sigma_t^2=\beta_{10}+\beta_{11}\epsilon_{t-1}^2+\beta_{12}\sigma_{t-1}^2$ and (2) an augmented GARCH model which is a no-jump version of Model 2 (i.e. $\lambda_t\equiv 0\ \forall t$) where $\sigma_t^2=\beta_{10}+\beta_{11}(1-RD_{t-1})\epsilon_{t-1}^2+\beta_{12}\sigma_{t-1}^2+\beta_{13}|PB_{t-1}|$. In the standard GARCH model $\beta_{11}+\beta_{12}=0.9977$ and in the augmented GARCH model $\beta_{11}+\beta_{12}=1.0244$, although the hypothesis that $\beta_{11}+\beta_{12}=1$ cannot be rejected by a likelihood ratio test at the 5% level of significance. Consistent with the work of Cai (1994) and Gray (1996a), allowing for jumps has dramatically decreased the persistence of shocks to exchange rates.[8]

4.2.3. Jumps

Next, we turn to the impact of jumps on the conditional distribution of exchange rates. Unfortunately, it is difficult to test for the absence of jumps, since under the null the parameters governing the jump are not identified (see Hansen, 1992 for a detailed discussion). In Table 6, the LRT statistic comparing our model with the nested no-jump model ($\lambda_t\equiv 0\forall t$), calculated in the standard manner, is 646.4021 which is exceptionally large by any benchmark. Despite the fact that we have not adjusted the χ^2 distribution of this LRT statistic to reflect the presence of parameters which are not identified under the null, we do gain some degree of confidence in the statistical significance of jumps from this exercise.

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[8] Bollerslev (1986) shows that a GARCH (1, 1) process can be written as an ARMA(1, 1) process in squared residuals, with autocorrelation parameter $\beta_{11}+\beta_{12}$. Whereas the half-life of a shock is more than 5 years in a standard GARCH (1, 1) model, in our model it is less than 2 weeks.

We can, however, formally test whether the jump probability is state-dependent or constant. The jump probability (λ_t) is negatively related to the slope of the yield curve in France ($\beta_2 < 0$)—when the yield curve inverts, the jump probability increases. Although the coefficient is not statistically significant (via the t-test in Table 5 or the LRT in Table 6), it is economically significant. What matters is the ability of the slope of the yield curve to predict large movements in exchange rates. For the majority of our sample, the jump probability does not show much variation, but spikes occur when speculative crises are expected. Hence, the identification of the coefficient comes primarily from those few observations where the yield curve experiences dramatic changes during speculative crises. Such a pattern can be captured by our model since the derivative of the jump probability with respect to the slope of the yield curve is steeply decreasing in its magnitude. In particular, when the yield curve is upward sloping, a 10% drop in the slope raises the jump probability by less than 0.025, but when the yield curve inverts the sensitivity to slope changes rises dramatically. Fig. 3 plots the jump probabilities which increase noticeably, prior to most realignments. The jump probability increases at other times to reflect the probability of non-realignment jumps. Once again, confidence bands are computed by generating 1000 samples of

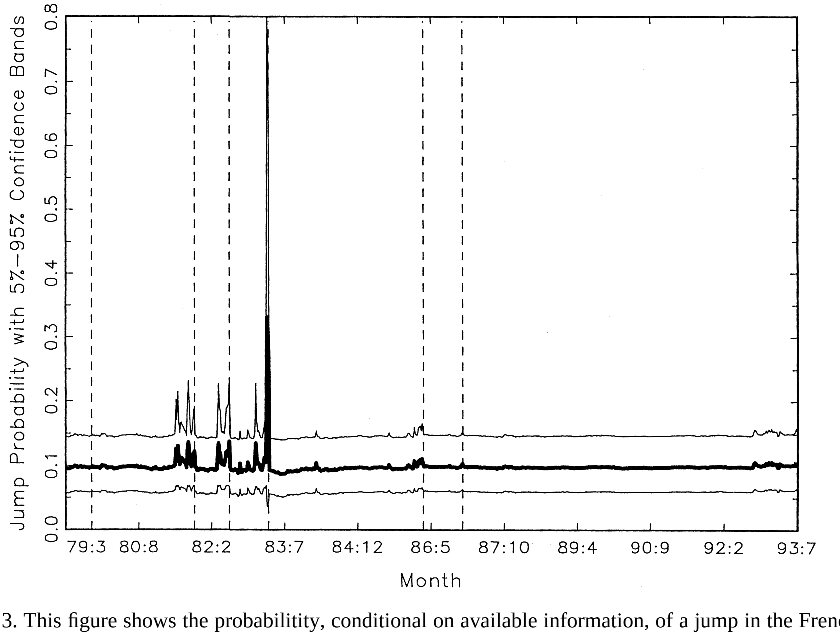

Fig. 3. This figure shows the probabilitity, conditional on available information, of a jump in the French Franc/Deutschmark (FF/DM) exchange rate in the following week (λ_{t-1}). The parameters of the model of the conditional distribution and the jump probabilities are estimated using the entire data set. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

parameter values from the multivariate asymptotic distribution of parameters. For each data point, the jump probability is recomputed for each set of parameters, yielding 1000 estimates of the jump probability for each observation. The confidence bands are then taken as the 5th and 95th percentile of the 1000 estimates.

The expected size of a jump, conditional on one occurring, (ρ_{t-1}) also varies substantially over time. A larger jump is expected when the level of reserves runs low ($\beta_4 < 0$), when the French interest rate differential with Germany increases ($\beta_6 > 0$) and when the cumulative inflation differential between the two countries is high ($\beta_7 > 0$). This indicates that weak-currency countries have their competitive position restored, at least partially, by realignments. Although the coefficient is not statistically significant, it is large in an economic sense. The effect of the position in the band on the expected jump size is not statistically significant. The time-variation in the expected jump size (ρ_t) is plotted in Fig. 4. Here the model indicates that the size of jumps is predictable. The expected mean jump size increases noticeably before the three large realignment jumps in the early 1980's and non-realignment jumps are expected to be of relatively smaller magnitude,

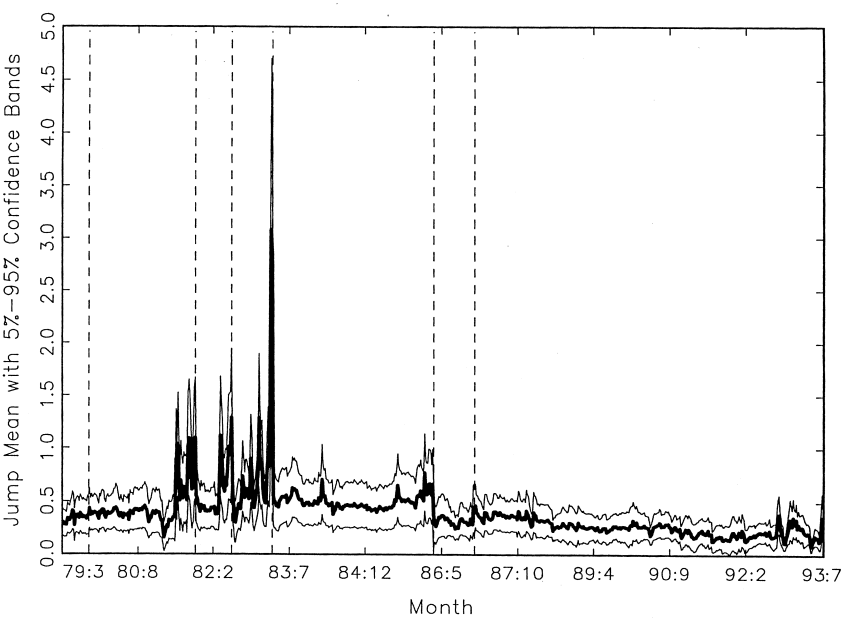

Fig. 4. This figure shows the expected size of a jump in the French Franc/Deutschmark (FF/DM) exchange rate in the following week, conditional on available information and on a jump occurring, (ρ_{t-1}). The parameters of the model of the conditional distribution and the jump means are estimated using the entire data set. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

which is also consistent with the observed data. Also, there is considerable movement in the confidence bands, consistent with significant time-variation in the jump size, which is confirmed by the joint significance of the conditioning variables, according to the LRT reported in Table 6.

Finally, we consider the impact of jumps on exchange rate volatility by introducing two measures of jump risk. First, note that the conditional variance of exchange rate changes can be written as:

$$\begin{aligned} h_{t-1} &= \mathrm{VAR}[\Delta S_t | I_{t-1}] \\ &= [(1-\lambda_{t-1})\sigma^2_{t-1} + \lambda_{t-1}\rho^2_{t-1}\delta^2] + (1-\lambda_{t-1})\lambda_{t-1}[\mu_{t-1} - \rho_{t-1}]^2 \end{aligned} \tag{9}$$

That is, jump risk consists of a variance term, $\lambda_{t-1}\rho^2_{t-1}\delta^2$, which is directly increasing in λ_{t-1} and ρ_{t-1}, and a conditional mean term. Since ρ_{t-1} can be significantly greater than μ_{t-1}, the latter can be quite important. We compute two measures of the relative importance of jump risk:

$$VR_{1_{t-1}} = 1 - \frac{(1-\lambda_{t-1})\sigma^2_{t-1}}{h_{t-1}}$$

and

$$VR_{2_{t-1}} = \frac{(1-\lambda_{t-1})\lambda_{t-1}[\mu_{t-1} - \rho_{t-1}]^2}{h_{t-1}}$$

From the results reported in Table 5, VR_2 is high when the mean jump size is higher than the expected exchange rate change within the band. Naturally, the ratio peaks before realignments but it never exceeds 25%. VR_1 adds a term that is increasing in the variance and probability of a jump. Interestingly, just before realignments virtually all of the exchange rate's conditional volatility is accounted for by jumps. Even in quiet periods, a significant portion of the total conditional volatility can be attributed to jump risk, in fact the sample mean of VR1 is 0.62. Note that jump risk was still substantial after 1987, a period often described as the most stable and credible period in EMS history. We discuss the credibility issue in more detail below.

4.3. Diagnostic tests

Finally, a series of diagnostic tests are performed in order to establish that the proposed model provides a reasonable description of the data. In particular, we compare the raw data, scaled to have zero mean and unit variance,

$$\Delta S^*_t = \frac{\Delta S_t - E[\Delta S_t]}{\sqrt{\mathrm{VAR}[\Delta S_t]}},$$

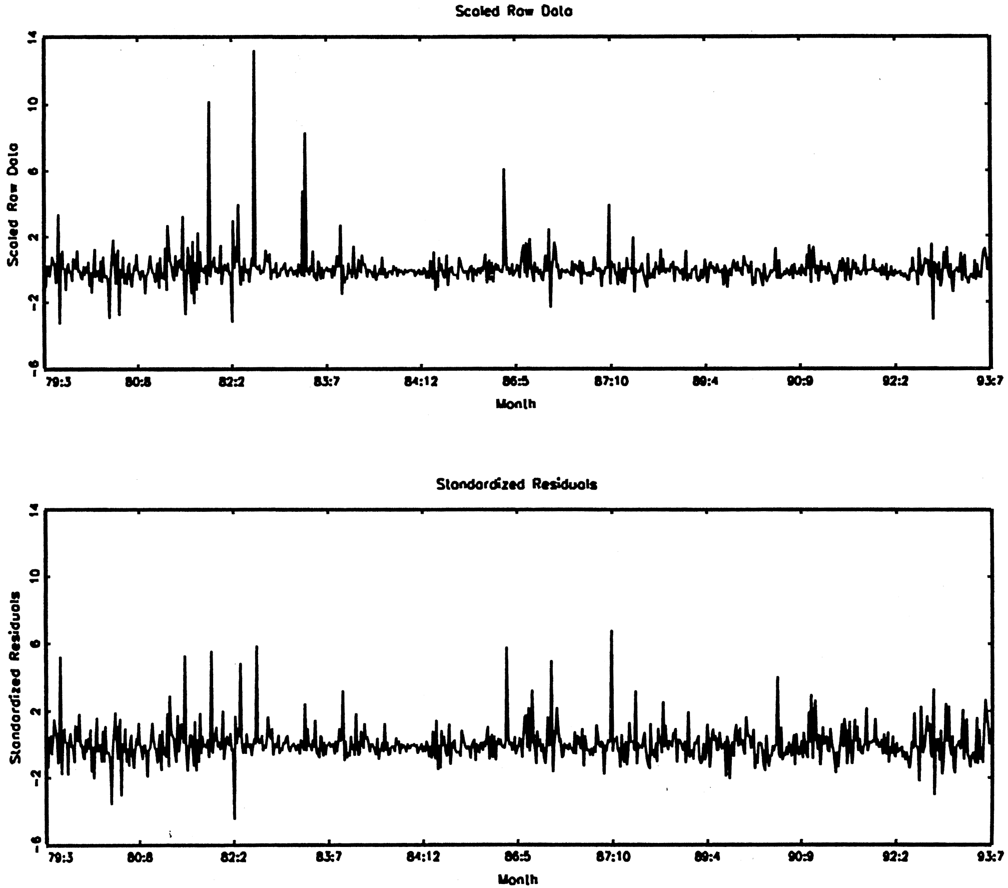

Fig. 5. This figure shows the scaled raw data $\Delta S_t^* = \frac{\Delta S_t - E[\Delta S_t]}{\sqrt{\mathrm{VAR}[\Delta S_t]}}$ and standardized residuals $z_t = \frac{\Delta S_t - E[\Delta S_t|I_{t-1}]}{\sqrt{\mathrm{VAR}[\Delta S_t|I_{t-1}]}}$. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

with the standardized residuals

$$z_t = \frac{\Delta S_t - E[\Delta S_t|I_{t-1}]}{\sqrt{\mathrm{VAR}[\Delta S_t|I_{t-1}]}}$$

These series are contrasted in Fig. 5. Whereas the raw data contains many outliers, the magnitude and frequency of outliers is much lower for the standardized residuals. Consistent with this, the kurtosis of the raw data is 64.64 whereas for the standardized residuals it is only 12.36. Moreover, the conditional heteroskedasticity that is evident in the raw data appears to be well-captured by the model. The p-values for the Ljung–Box statistics for serial correlation in the absolute value of the raw data are 0.0330, 0.0030, 0.0084, 0.0138, and 0.0280 for the first five lags, indicating significant serial correlation. For the standardized residuals the corresponding p-values are 0.5939, 0.8399, 0.8012, 0.8527, and 0.6834, indicating that the autoregressive volatility is well captured by the model.

5. The credibility of target zones

5.1. Credibility and realignment probabilities

Several authors have studied the credibility of the EMS. Rose, Svensson (1993) Rose, Svensson (1995), Frankel, Phillips (1992) and Chen, Giovannini (1992) rely on (1) uncovered interest rate parity and (2) an estimate of the expected exchange rate change within the band, to infer realignment expectations. The consensus in the literature seems to be that the EMS was essentially credible from 1987 to 1991—a period that contains no realignments. Consequently, the crises in September 1992 following the signing of the Maastricht Treaty in December 1991, and the eventual breakdown of the system came as a surprise. In this section, we examine whether our model implies a similar increase in credibility after 1987.

Recall that we define a target zone to be ‘‘perfectly credible’’ if there is zero probability that it will be realigned. Let p_{t-1} denote the realignment probability at time $t-1$ which is $Pr[S_t > \ln(U_{t-1})|I_{t-1}]$. This is the probability, conditional on available information, that the exchange rate will move above the upper boundary of the target zone. Since our model fully specifies the conditional distribution of the exchange rate, we can easily compute this conditional probability. Interestingly, unlike the papers listed above, we can also disentangle the magnitude and probability of a realignment. For example, we can compute the probability that next week’s exchange rate will be 5% above the current upper band.

The realignment probabilities, p_{t-1}, are plotted in Fig. 6. The realignment probability spikes noticeably before all six realignments. Interestingly, the mean realignment probability appears to be higher in the period since 1988. We interpret this as evidence of the predictability of the eventual breakdown of the system. That

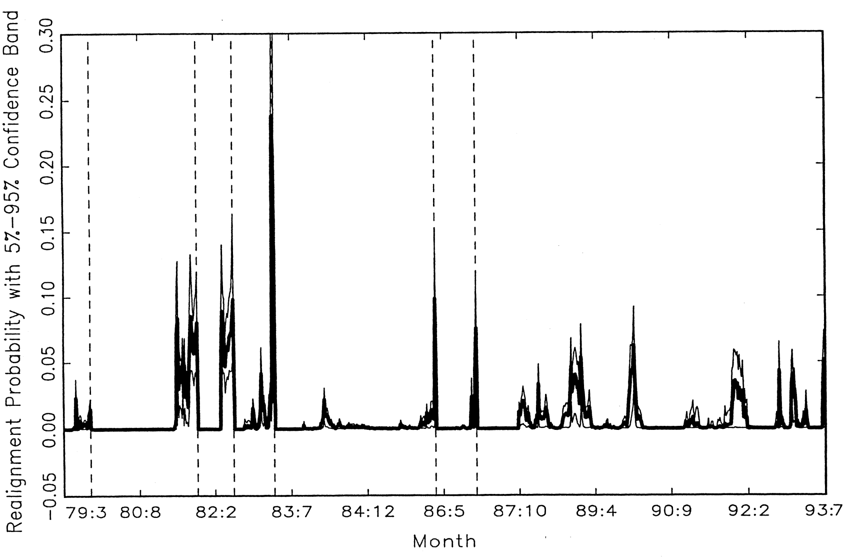

Fig. 6. This figure shows the probabilitity, conditional on available information, that the French Franc/Deutschmark (FF/DM) exchange rate will move outside the European Monetary System target zone during the following week. This conditional probability is computed by integrating the area of the conditional distribution which is outside the target zone, and is interpreted as the realignment probability. The parameters of the model of the conditional distribution and the realignment probabilities are estimated using the entire data set. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

is, in the post-1988 period, conditions which would previously have caused a realignment, caused no change to the system. This resulted in a sustained build up of realignment pressure. When no pressure-relieving realignments were forthcoming, the entire system eventually broke down. The confidence bands are relatively tight around the realignment probability and vary considerably over time, consistent with significant time-variation in the conditional probability of a realignment.

The two potential criticisms of this argument, are addressed in turn below. First, our analysis is within-sample and the apparent realignment predictability is potentially spurious. In the next sub-section, we conduct a true out-of-sample analysis of the predictive power of the model. Second, structural changes, including changes in capital controls and the signing of the Maastricht Treaty, may have affected the structural parameters of the unspecified underlying model. Therefore, our reduced form parameter estimates may be unstable.

5.2. Predictability of realignments

Although the variables that we use in predicting realignments have been used before, we argue that our analysis does not amount to data-snooping because most previous work has not found evidence that these variables hold any predictive power. Moreover, we are also careful to allay concerns about some form of ‘‘model-snooping’’. That is, the model specified in Eq. (3) relies on a thorough analysis of how the EMS operated during the sample period. In 1979, it is unlikely that somebody could have formulated such a model. For example, the EMS had a number of mechanisms (such as the divergence indicator) designed to make it a symmetric system, rather than the DM-anchored system that it eventually became. Consequently, we attempt to construct a series of true ex-ante realignment probabilities, noting that our econometric model could only have been formulated after a number of realignments had occurred, clarifying the role of the DM, the effect of speculative crises on interest rates, and the importance of real exchange rate changes. Therefore, we use data from 1979 to 1983 (the first 199 observations) and focus on the predictive power of the model for further realignments (the remaining 550 observations). To do so, we re-estimate the model every week adding new observations and use these parameters to compute the realignment probability one week ahead, conditional on observable information. These probabilities are plotted in Fig. 7.

The annualized root mean square error (RMSE) from using our model to predict the exchange rate one week ahead is 15.5% compared to 16.2% and 16.5% for the RMSE of a simple random walk model $E[S_{t+1}|I_t]=S_t$ and the unbiasedness model $E[S_{t+1}|I_t]=S_t+(i_t^{\mathrm{fu}}-i_t^{\mathrm{gu}})/52$, respectively. Whereas this improvement in out-of-sample fit may be small, Meese, Rogoff (1983) note that most empirical and structural exchange rate models fail to beat the random walk model. Moreover, we repeated this experiment using a more parsimonious model with $\beta_5=\beta_7=\beta_9=\beta_{13}=0$. This model yields a RMSE of 15.1%. Finally, the annualized RMSE of the 36 largest exchange rate changes (5% of the sample) are 53.2%, 53.0%, and 47.3% for the random walk, unbiasedness model, and our model, respectively. We draw some comfort from these results since a reliable forecast of the largest exchange rate changes is particularly important for most market participants.

5.3. Credibility and structural changes

Controls on international capital flows can protect domestic interest rates from the large fluctuations associated with expectations of exchange rate realignments. Moreover, a central bank may experience large losses of reserves when holders of domestic high-powered money sell the domestic currency to the central bank in exchange for foreign currency just before a devaluation is expected. To some, capital controls were an essential ingredient of the stability of the EMS (see, for

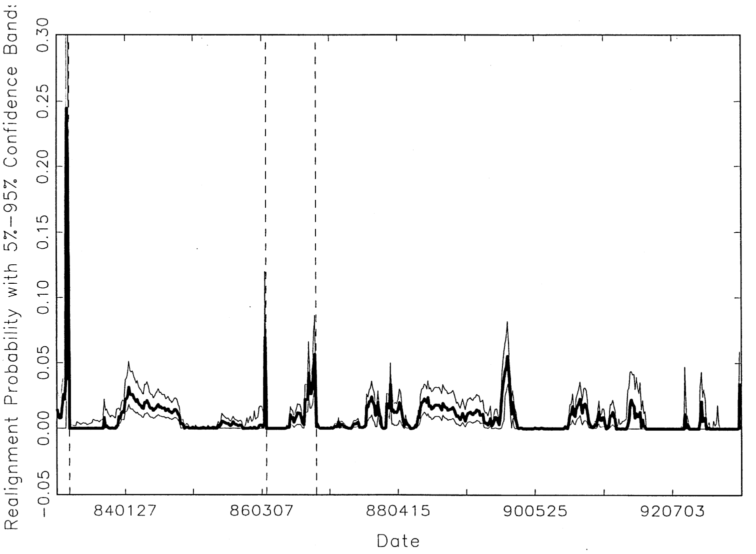

Fig. 7. This figure shows the probabilitity, conditional on available information, that the French Franc/Deutschmark (FF/DM) exchange rate will move outside the European Monetary System target zone during the following week. This conditional probability is computed by integrating the area of the conditional distribution which is outside the target zone, and is interpreted as the realignment probability. In computing the realignment probability at time t, the parameters of the model of the conditional distribution are estimated using data up to time $t-1$ only. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

example, the famous Padoa-Schioppa report of Padoa-Schioppa, 1985), to others, capital controls prevented optimal allocation of resources and may have been largely circumvented anyway.

France has a long history of capital controls, but they were gradually relaxed during our sample period, especially after 1986.[9] If changes in capital controls affected the credibility of the system, their effects would be most visible in the parameters governing the size and probability of a jump. Unfortunately, this makes it virtually impossible to construct meaningful tests for structural change. Since most changes occurred after 1985, but no major realignments occurred during this period, it is impossible to identify the jump parameters using only post-1985 data. When estimating realignment probabilities using only post-1985 data, one is bound to find low realignment probabilities. However, although no major realignments occurred, this need not mean agents did not expect them. That is, the post-1985

[9] A previous version of this paper, Bekaert, Gray (1996), provides a survey of these capital controls and when they were relaxed.

data potentially suffer from a classic peso problem. Consequently, it is critical to use the early, more turbulent, period to estimate the jump parameters.

Although this prevents us from conducting formal econometric tests, we can use Fig. 7 to study the behavior of the out-of-sample realignment probabilities around the dates of potential structural changes. The first issue is whether the relaxation of capital controls affected the credibility of the system. Many doubted the sustainability of the EMS in light of the classic argument of the impossibility of pursuing an independent monetary policy in a system of fixed exchange rates and perfect capital mobility. However, most empirical studies find that the credibility of the EMS increased considerably after 1987 (see Frankel, Phillips, 1992; Chen, Giovannini, 1992, and Rose, Svensson, 1993). Hence, these studies suggest that capital controls are not a necessary condition for a credible EMS. The currency crises in 1992, however, and the ensuing reimposition of capital controls by Spain, casts doubt on this conclusion.[10] Our results have somewhat different implications. As Fig. 7 shows, realignment probabilities remain high after 1987 and there is no downward trend. This is all the more surprising, since inflation differentials, one of the underlying macroeconomic causes of tensions, had substantially narrowed over time. Interestingly, one period of relative turbulence occurs shortly after the June 1988 EC decision to remove all capital controls in the EMS. This indicates that the post-1987 period may have been less stable than previously thought.

The second issue is the impact on credibility of the Maastricht Treaty, which laid out the process for economic and monetary union (EMU). While there is an extensive debate about the costs and benefits of EMU, the question is ultimately an empirical one. In the context of our model, if the Maastricht Treaty was fully credible, the realignment probability should have decreased after (or even before) the Treaty was signed.

Fig. 7 shows that there is a clear increase in realignment probabilities in November 1991, which is reversed by January 1992. This may reflect the speculation of market participants about the possibility of one last realignment before the process towards EMU was started, or general uncertainty about the feasibility of the Treaty. In fact, after German reunification in 1990 many economists argued that a revaluation of the DM was warranted and that failure to do so may put strains on the movement towards EMU. In contrast to previous studies (see especially Rose, Svensson, 1993), we find evidence of such strains. For example, after the fall of the Berlin Wall in November 1989 and the December 1989 Strasbourg Summit, in which a date for the Maastricht Conference was agreed upon, realignment probabilities increased significantly reaching 5.5% at year-end.

[10] Interestingly, Ireland endured a 10% devaluation in January 1993, one month after its capital controls were lifted, despite having strong fundamentals. The exit of the British pound was a major factor in the pressure on the Irish punt, but it is striking that the devaluation could no longer be averted once capital controls were lifted.

The currency crises have been the subject of even more debate, particularly the September 1992 crisis which led to both the British pound and Italian lira exiting the system. Was the crisis anticipated by the markets? It is interesting to re-examine this question in the context of our model, since Rose, Svensson (1993) and Rose (1993) have concluded from a quite different framework that the currency crisis in September 1992 came as a surprise to market participants and governments alike.[11]

Surprisingly, there is no marked increase in realignment probabilities during 1992 and our model does not show any effect of the rejection of the Maastricht Treaty by the Danes on the credibility of the FF/DM band. One week before the French referendum on September 20, there was an increase in the realignment probability. The turbulent period afterwards with devaluations of the peseta and the escudo, and the suspensions of the ECU links by Sweden and Norway, generated little loss of credibility for the French Franc, except in December 1992. In 1993, realignment probabilities are close to zero, consistent with the credibility of the FF/DM target zone. This coincides with a period in which French short and long-term interest differentials virtually converged to German levels, and some market observers talked about the ''franchor'', the French Franc replacing the DM in the anchor role of the EMS. There is a marked increase in the realignment probability in the week of July 23, 1993, which is large by historical standards. This is the week prior to the August 2 crisis when the parity bands were widened to 15% on either side of the central rate. Hence, our model would have produced a useful warning signal of the trouble ahead.

In summary, we find that 1992 was not a ''turbulent'' year, relative to historical averages. This may indeed mean, as Rose, Svensson (1993) argue, that ''the currency crisis may have been caused by phenomena without long gestation lags of the sort that characterize most macroeconomic and political events.''

One structural change does not suffer from a peso problem and can be formally tested. The Basle–Nyborg Agreement of September 1987 intended to strengthen the ERM by providing for intra-marginal intervention and more liberal short-term financing of interventions. This agreement may affect the way intra-marginal intervention is conducted, potentially affecting the reversion of exchange rates towards the center of the band. We formally test this by allowing the conditional mean parameters, β_8 and β_9, to take different values before and after the agreement. Neither of these extra parameters is individually significant and the LRT of their joint significance, which is χ^2_2 under the null, is 0.5986 which is not significant at any usual level. We conclude that the Basle–Nyborg Agreement merely formalized the practice of intramarginal intervention which was already

[11] One disadvantage of our framework is that our model does not specify the full dynamics of all variables used to predict exchange rates. Hence, we can only look one week ahead and cannot sketch the evolution of longer-term expectations.

common in the EMS. This empirical result is also consistent with the theoretical analysis in Flandreau (1996) and Serrat (1995) showing how intramarginal intervention arises endogenously in the context of a multilateral target zone model.

6. Implied foreign exchange risk premia

Svensson (1992a) argues that the foreign exchange risk premium in a target zone is small. His analysis is based on a simple optimizing model with exchange rate uncertainty arising from exchange rate movements inside the band and occasional realignments which follow a Poisson process. He concludes that risk premia arising from exchange rate movements within a narrow band, as are in place in the ERM, are insignificant whereas risk premia arising from devaluation risks may be considerably larger but are still relatively small in comparison with the expected rate of devaluation. His results are important because they have motivated a large literature on the computation of realignment probabilities (see above), which imposes UIRP and ignores risk premia.

The model developed in this paper provides a challenge for these results and the methodologies on which they are based. Since we model the complete conditional distribution of exchange rate changes, we can compute the implied risk premium on FF investments for German investors as $i_t^{\mathrm{fu}} - i_t^{\mathrm{gu}} - 52E[\Delta S_{t+1}|I_t]$. In annualized percentage terms, the risk premium has a mean of 3.02% with a standard deviation of 3.28% and a first order autocorrelation coefficient of 0.867. In contrast to a maximum risk premium of 4.5% in Svensson's model, the risk premium graphed in Fig. 8 varies between -3.30% and 31.45%. The risk premium seems to satisfy the Svensson bands most of the time, but increases substantially in times of speculative crises before realignments. This suggests that the jump risk discussed above is priced. Regressions of the risk premium on the variance ratios VR_{1_t} and VR_{2_t}, which measure the importance of jump risk, yield highly significant positive slope coefficients.[12] When speculative crises hit, both the expected rate of devaluation and the uncertainty about future exchange rate movements increase dramatically. However, whereas Svensson (1992b) claims that the resulting increase in the interest differential between French and German deposits is primarily due to the increase in the expected rate of devaluation, we find that a substantial part of the increase in the interest differential reflects currency risk.

This debate parallels the debate on the size of the foreign exchange risk premium for floating currencies. One interpretation of the empirical evidence on UIRP, implies the existence of highly variable risk premia. For example, Bekaert (1995) uses a vector autoregressive framework to empirically derive lower bounds

[12] Because both the risk premium and the variance ratios contain measurement errors that may be correlated, this analysis is only suggestive.

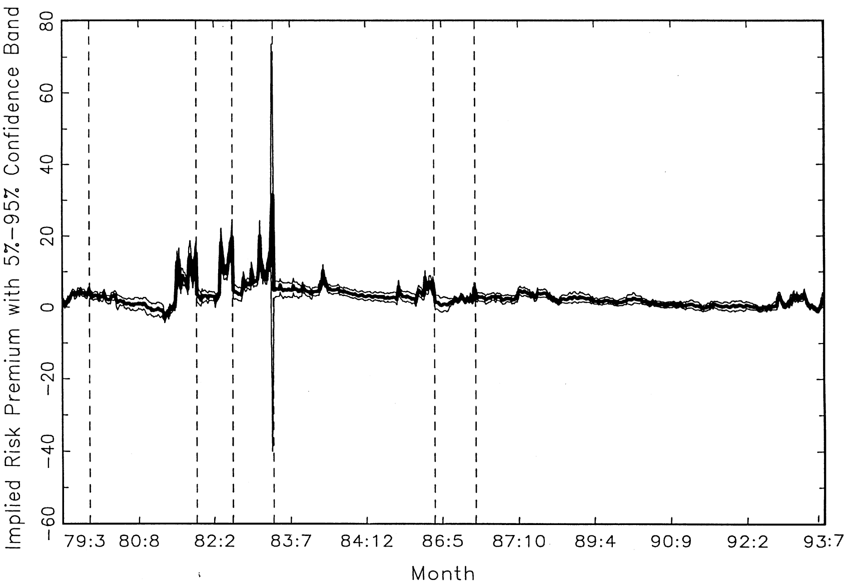

Fig. 8. This figure shows the implied risk premium for the French Franc/Deutschmark (FF/DM) exchange rate. This is computed as the difference between the interest differential and the expected exchange rate change, and is reported in annualized percentage terms. The parameters of the model of the conditional distribution and the jump means are estimated using the entire data set. The dashed lines indicate realignments of the FF/DM target zone. The data set consists of weekly observations from 23 March 1973 to 23 July 1993, a total of 749 observations.

on the variability of risk premia on yen, pound and mark investments for U.S. investors (and all cross-rate investments). He finds the variability of these premia to be in the order of 10%, three times as large as our estimate for the FF/DM risk premium. Moreover, the risk premium changes sign and is often quite large. Although here too it is often claimed that the risk premium is small (see Frankel, 1988), these claims are always model-based. Many fundamentals simply do not show the required variability to explain the empirical evidence on UIRP deviations.[13] Similarly, it is not surprising that Svensson's theoretical model fails to generate the required variability. For example, he assumes that the exchange rate volatility within the band only depends on the position in the band as in the Krugman (1991) model. We have shown above that volatility within the band exhibits marked GARCH effects and that the dependence on the band is contrary to what is predicted by the Krugman model.

[13]See Bekaert (1994) for the analysis of the foreign exchange risk premium in a monetary general equilibrium model.

While in the presence of unsatisfactory theoretical models, one may be tempted to rely on empirical estimates of the risk premium, there are dangers in this approach as well. The reduced-form estimation of the conditional distribution of the exchange rate is subject to small sample problems and the existence of peso problems may generally make it difficult to infer the correct probability distribution actually used by agents from a finite data set. Nevertheless, our empirical results and the out-of-sample analysis discussed above instill some degree of confidence in our risk premium estimates. Moreover, there are few alternative empirical approaches. One promising alternative approach is to use options data to infer the exchange rate's conditional distribution as in Campa, Chang (1996).

Our findings have a number of implications. First, the variability of risk premia within a target zone is considerably smaller than empirical estimates of the variability of risk premia for floating exchange rates. Second, the risk premia are sizable and should not be ignored. Hence, the practice in the recent target zone literature (e.g., Rose, Svensson, 1993, Chen, Giovannini, 1992) of relying on UIRP and disregarding the risk premium may yield unreliable empirical estimates of realignment probabilities.

7. Conclusions and future work

This paper develops a rich empirical model of exchange rates in a target zone and applies it to the FF/DM rate.[14] In contrast to some recent empirical analyses, we detect substantial non-linearities in the behavior of the FF/DM rate.[15] We also find that, in addition to realignments of the target zone itself, exchange rates exhibit a tendency to jump within the target zone. Our model is able to predict the likelihood and size of these jumps. Furthermore, by modeling the entire conditional distribution of exchange rates, we are able to isolate the probability of target zone realignments. In contrast to previous work, we show that realignments and the eventual breakdown of the system are predictable and that the credibility of the EMS did not increase after 1987. Moreover, the popular practice of computing realignment probabilities imposing UIRP ignores the potentially important impact of large foreign exchange risk premia before realignments.

Our work has implications that extend beyond the realm of the EMS. The recent proposals to limit the variability of floating exchange rates by means of target zones implicitly assume that target zones effectively reduce the variability of exchange rates, and hence limit the costs of exchange rate uncertainty. Our findings indicate that when a target zone is imperfectly credible, exchange rate

[14] In Bekaert, Gray (1996), we apply a modified version of this target zone model to the Deutschmark/US Dollar rate and find no evidence of target zone behavior.

[15] Gourinchas (1995) reaches similar conclusions using a non-parametric instrumental variables approach.

variability can remain substantial because of the presence of jump risk. Moreover, this risk seems to be priced leading to enormous yield differentials between currencies before a realignment is expected. Even barring arguments about the true costs of exchange rate uncertainty, it is not clear to us that replacing a system of floating exchange rates, exhibiting high variability, with a target zone system, with lower variability on average, but occasional extreme volatility, is welfare-improving. In particular, risk stemming from a volatile diffusion is much easier to hedge than that stemming from a jump process, even if the volatility between jumps is relatively low.

There exist a number of avenues for future research which could be undertaken to further flesh out the behavior of exchange rates within a target zone. First, we have only considered a bilateral exchange rate in isolation. In a system such as the ERM, movements in third currencies can put pressure on the FF/DM rate. That is, there are effective bilateral bands which are narrower than the actual band (see Pill (1994), Flandreau (1996), and Serrat (1995)). Unfortunately, it is likely to prove numerically infeasible to extend our techniques to an entire system of exchange rates. In this sense, the realignment probabilities we compute may under-estimate the true realignment probabilities. Second, a bivariate model of exchange rates and interest rates could be developed in order to examine UIRP in the context of our dynamic setting. Dahlquist, Gray (1996) demonstrate that there is evidence of regime-switching behavior in EMS short-term interest rates. In a credible target zone, it is problematic to test unbiasedness using a linear regression since the interest differential is likely to be correlated with the error term. Moreover, the possibility of infrequent but large realignments makes the EMS an ideal laboratory for the analysis of peso problems.

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