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Abstract

In this paper we reconsider the relationship between spot and forward rates, augmented by a term which contains a measure of conditional volatility. Previous parametric specifications such as the GARCH-M provided disappointing results possibly due to the degree of persistence on the estimated conditional volatility. Instead, we propose a semiparametric estimator based on a nonparametric measure of the conditional volatility and we estimate the relationship with monthly data on six currencies *vis–à–vis* the Deutsche Mark . Another advantage of such a procedure is that data available at different frequencies can be used, as well as an indicator of market sentiment in the form of trading signals to purchase or sell a currency.

Keywords: Exchange Rates; Risk–Premium; GARCH-M; Semiparametric Estimation; Technical Analysis.

1 Introduction

The comparison between exchange rate movements and interest rate differentials shows that the profile of the former is too much complex to be explained in terms of the latter alone. In fact, the statistical tests generally fail to support the hypothesis that the interest rate differentials are an unbiased predictor of exchange rate movements (cf. Baillie and McMahon 1989). Among the explanations extensively studied in the literature, alternative theories have been proposed which call into discussion the issues of market efficiency, rational behavior, presence of a *peso problem*, and possible nonlinear dynamics being generated on the markets by the presence of heterogeneous agents.

Among the theoretical suggestions, the Lucas (1982) model of intertemporal asset pricing in a two–country world was adapted to show (e.g. Hodrick and Srivastava 1984) that uncertainty about the future purchasing power of domestic and foreign monies, and about future marginal utility of the domestic good translates into uncertainty about the intertemporal rate of substitution of domestic currency between t and a future date $t+k$. The presence of a conditional covariance term between this rate of substitution and the future spot rate is used to support the argument for the existence of a time–varying risk–premium.

One of the difficulties with this theoretical model is that the hypotheses entailed by it are not testable without paying the price of inserting strong assumptions in order to derive an estimable relationship. For example, a proposed model for the spot/forward exchange rate relationship considers a measure of conditional volatility in the equation for the mean. The result is a statistical model with the goal of extracting an economically interpretable signal from the excess returns on forward positions. One puzzling aspect of this approach seems to be that conditional volatility is always positive irrespective of the chosen *numéraire*, and does not refer to a risk-free asset as in the case of the CAPM model, challenging the possibility of giving it a risk–premium interpretation. A monotonic and increasing relationship between conditional variance and risk–premium has been recently challenged by Backus and Gregory (1993) who claim that the use of the conditional variance as a proxy for the risk–premium can be justified on the basis of a specific structure of the economy, but is by no means general. In what follows, we will show how the *impact* of conditional volatility can indeed be positive or negative once additional information on the exchange rate behavior in different market situations is taken into account.

Domowitz and Hakkio (1985) have exploited the parametric ARCH-M specification in which the equation for the mean in the spot/forward relationship is augmented by a conditional variance term which is assumed to follow an ARCH process. Their evidence (on monthly data) fails to give strong support to the existence of a risk–related effect on exchange rate movements. In this paper we will re-examine the reasons for this failure using monthly data ourselves providing an explanation for the disappointing performance of their model, in terms of the estimated persistence in the variance. The change in the size of the risk premium and the frequent changes in sign discovered by Stockman (1978) were interpreted then as being related to the nature of the stochastic processes ruling the state variables. Adding to that the highly nonlinear nature of the transformations these processes undergo in the intertemporal asset pricing models, the adoption of a nonparametric measure of risk seems to buy a lot of flexibility relative to a parametric specification. For this reason we propose a semiparametric estimator (derived from Generalized Method of Moments conditions) which exploits the local approximation properties of kernel estimators. Since the estimated conditional variance is based on the residuals of an auxiliary regression, one advantage is to reduce the degree of persistence. Also we have the possibility of using higher frequency data (say, weekly) in deriving the instruments for estimating the impact of the conditional volatility. A comparison between the evidence produced with the instruments estimated on monthly data and the ones estimated on weekly data (and then sampled at a monthly frequency) shows that the latter provides a more significant impact of the risk–premium term on exchange rate movements.

In this context, the need for the measured conditional variance to reflect as much as possible recent market conditions is made even clearer when we consider the additional information provided by trading signals suggested by technical analysis. We label periods as "buy" or "sell" on the basis of the joint outcome of simple trading rules, and then we analyze the differentiated impact of the conditional volatility on exchange rate movements. We take the trading signals as an indicator of the market sentiment about the direction of a currency. In doing this we feel comforted by the high diffusion of technical analysis tools among traders reported by Taylor and Allen (1992) and by the results obtained by LeBaron (1993a,b) on the possibility of detecting profitability from adopting the signals as a trading strategy.

We show that the impact of conditional volatility on exchange rate movements changes sign across "buy" or "sell" periods and therefore gives a different interpretation to the question of time-varying/sign-changing risk–premium. Therefore, we are capable of motivating the interest in the impact of the conditional variance since it measures the perceived risk in detaining a currency when the currency is appreciating ("buy") or depreciating ("sell") *vis–à–vis* the numéraire. In a way, the approach is similar in spirit to the analysis by Engel and Hamilton (1990) who try to discover whether information about "long swings" in the exchange rates (long periods of appreciation or depreciation detected through a Markov-switching model) make a difference for the uncovered interest rate parity (UIRP) relationship. Also in the present case the answer is negative, in that UIRP is rejected: on the positive side, though, our suggestion provides a tool by which a "fad" on the market (a run on buying or selling a currency) can be assessed and its importance evaluated.

The structure of the paper is as follows: in section 2 we set the notation and discuss the risk–augmented specification in its parametric (G)ARCH-M form. In section 3 we propose the GMM-based semi-parametric estimator. In section 4 we highlight the empirical limits of a parametric specification with reference to six currencies *vis–à–vis* the Deutsche Mark (French Franc, Italian Lira, British Pound, Japanese Yen, Canadian Dollar, and US Dollar) with monthly data from June 1973 to

January 1994 (248 observations). In section 5 we propose the empirical evidence for the six currencies from our monthly specification when the the impact of conditional variance is estimated either with monthly-based or with weekly-based instruments. Finally, the characteristics of "buy" and "sell" periods and the time profile of the impact of the conditional volatility on exchange rate movements are presented in section 6 where historical evidence for asymmetric effects is produced. Concluding remarks follow.

2 The Model

The theory of interest rate parity states that, in the absence of market frictions, transaction costs, capital controls, and so on, when faced with the need of availability of foreign currency k periods into the future one would be indifferent (in *ex ante* expected terms) between holding domestic currency (lucrating domestic interest rates) and purchasing a forward contract or purchasing foreign currency (lucrating foreign interest rates) right away.

In particular, uncovered interest rate parity (UIRP) is at the basis of many econometric models, although it has received very little empirical support in practice. The relevant relationship can be written as:

$$E_t(s_{t+k}) - s_t = \log(1 + i^d_{t,k}) - \log(1 + i^f_{t,k}) \approx i^d_{t,k} - i^f_{t,k} \equiv i^*_{t,k}. \quad (1)$$

where s_t is the (logarithm) of the spot rate at time t expressed as units of foreign currency per unit of domestic currency; E_t is the expected value conditional on the relevant information set at time t, $i^d_{t,k}$ is the interest rate on the domestic currency between t and $t+k$; $i^f_{t,k}$ is the interest rate on the foreign currency on the same horizon and on foreign assets perfectly substitutable with domestic ones. Consider that the interest rate differential is used to form the so-called forward premium (covered interest rate parity), $f_{t,k} = s_t + i^*_{t,k}$, where $f_{t,k}$ is the (logarithm) of the forward exchange rate at time t for delivery at time $t+k$, and therefore provides the basis on which forward rates are determined in practice from spot rates. UIRP then becomes

$$E_t(s_{t+k}) - s_t = f_{t,k} - s_t.$$

Such a relationship is a convenient one because it avoids problems connected to the possible nonstationarity of the exchange rate series when converted into an estimable/testable form

$$s_{t+k} - s_t = \beta_0 + \beta_1 \ (f_{t,k} - s_t) + u_{t+k}. \tag{2}$$

As discussed by Froot and Thaler (1990), the issues of whether the difference $s_{t+k} - f_{t,k}$ (under the hypothesis $(0,1)$ for (β_0, β_1) from (2)) has a zero mean (unbiasedness hypothesis), is uncorrelated, or has a constant variance have received considerable attention in the literature obtaining results which vary a lot among themselves, according to which currency was under consideration and for what period.

When considering the graphical evidence of the behavior of the two variables with respect to time (reported in the top panel of Figure 1 for the French Franc/Deutsche Mark exchange rate as an example) one can see that the signal contained in the forward premium $f_{t,k} - s_t$ is smooth relative to the dynamics exhibited by the exchange rate movements $s_{t+k} - s_t$. In particular, a fairly stable interest rate differential (positive for the French Franc for most of the period under exam) is accompanied by wide swings in the exchange rate showing that other elements are at work and should be investigated. For the exchange rate at hand, in particular, we inserted vertical bars in correspondence to the inception of the Exchange Rate Mechanism of the European Monetary System (March 1979) and to realignments of central parities. The date of the exit from the ERM by the Lira and the Pound (September 1992) is reported for the French Franc since it marked a period of crisis for that currency as well.

A similar picture emerges from a different representation of the same data in a cross-scatterplot (bottom panel of Figure 1), where the points corresponding to the realignment dates are marked with a cross. Somewhat surprisingly, some of the crosses lie deeply around the origin of the plot suggesting that there was no message about the possibility of a realignment in the interest rate differentials *at the end of the month* prior to the realignment. This is due to the speed at which the crises have occurred, with interest rate differentials exploding just for (at most) a few days prior to the establishment of a new central parity. Since the behavior of the exchange rate is constrained by institutional mechanisms within the ERM, the issue was raised by several Authors (e.g. Svensson 1993) about the presence of a dichotomy between expected rate of change *within* the band and *without* which should be considered when evaluating the ERM credibility. In that stream of literature, interest rate differentials are taken to be a good approximation to the expected rate of change in exchange rates. In this paper we take quite a different stance, arguing that the information contained in the interest rate differential in itself does not convey enough message about the perception of the risk involved in detaining a certain currency and can be supplemented by other elements which play a more important role in practice.

In what follows, we will propose a different way of processing information, pursuing the argument which focuses on the effects of higher moments on the mean, relating what is left in the exchange rate movements (after the forward premium is taken into account) to a risk interpretation.

Although the links with economic theory are tenuous, the risk–related explanation of why (2) does not hold has received empirical attention since the seminal paper by Stockman (1978), who pointed out the presence of a time–varying risk premium, and the frequent change in sign when the estimation period was divided up into sub-samples. The lack of a theoretical model which can be translated into empirical testability is at the basis of the various statistical models of risk where the goal of the analysis becomes one of extracting an economically interpretable signal from u_{t+k}.

In fact, in order to investigate the relevance of the risk–related argument let us keep separate within u_{t+k} two terms, one which we will label $RP_{t,k}$ representing the risk–premium of the theory and the other a random disturbance ϵ_{t+k}. We have the expression

$$s_{t+k} - s_t = RP_{t,k} + \beta_1(f_{t,k} - s_t) + \epsilon_{t+k}. \qquad (3)$$

$RP_{t,k}$ is assumed to be linked to the conditional variance in the ϵ_{t+k}.

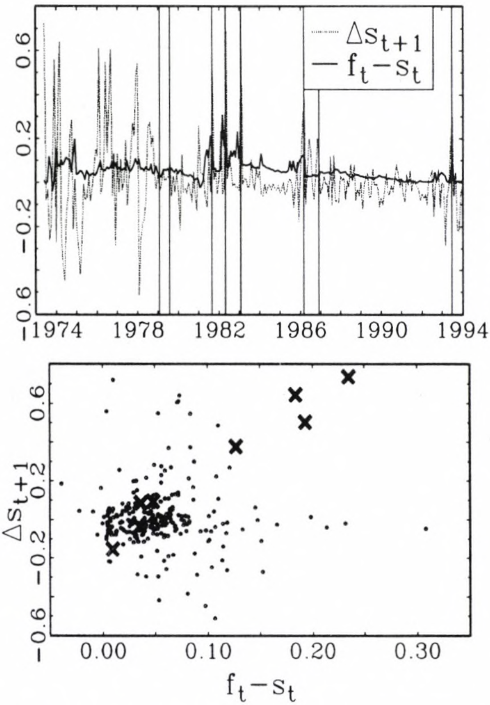

Figure 1: French Franc: Exchange Rate Movements and Forward Premia 1973-1994.

3 A Generalized Method of Moments Approach

Let us rewrite (3) considering, for the sake of simplicity, the case of one-month maturity on forward contracts and monthly growth rates for the exchange rates (to simplify notation f_t will stand for $f_{t,1}$) and let us consider an unobservable volatility term $\sigma^2_{t+1|t}$ conditional on the information set Ψ_t to be inserted in the spot/ forward relationship

$$\Delta s_{t+1} = \beta_0 + \beta_1(f_t - s_t) + \delta\sigma^2_{t+1|t} + \epsilon_{t+1}. \tag{4}$$

This model requires a special treatment from an econometric point of view. Among the solutions suggested, the ARCH-M model employed by Domowitz and Hakkio (1985) consists in adopting a parametric specification for the conditional variance. Following Engle, Lilien and Robins (1987), the risk term $RP_{t,1}$ is specified as being a linear function of the conditional variance of the error term of the type

$$RP_{t,1} = \beta_0 + \delta h_{t+1}$$

where h_{t+1} is defined from the conditional distribution of

$$\epsilon_{t+1} \mid \Psi_t \sim N(0, h_{t+1})$$

and follows a general ARCH(p) representation as

$$h_{t+1} = \alpha_0 + \sum_{i=1}^{p} \alpha_i \epsilon^2_{t+1-i} + z_t\phi,$$

Ψ_t is the information set available at time t and z_t is a vector of variables belonging to the information set of interest for the analysis (for example, dummies). In such a model the conditional variance is evolving as a function of its own past and enters the equation for the mean as well through $RP_{t,1}$. In the empirical section we will also consider the GARCH-M model as an extension in which the conditional variance can be expressed as

$$h_{t+1} = \alpha_0 + \sum_{i=1}^{q} \alpha_i \epsilon^2_{t+1-i} + \sum_{j=1}^{p} \gamma_j h_{t+1-j} + z_t\phi.$$

By its own nature, this term is time–varying and lends itself to act as a risk term once the signs of β_0 and δ are determined. Remark that opposite signs of β_0 and δ are capable of providing a risk–premium which would switch sign through time as a consequence of the size of the estimated conditional volatility.

The disappointing results of the analysis by Domowitz and Hakkio (failing to lend support to the importance of conditional variance in the explanation of exchange rate movements) have been attributed to the use of monthly data; other authors think that the univariate framework is too restrictive, while in a multivariate framework one could take into consideration not only the conditional variances but also the covariances among the various currencies in the market. Yet, Bollerslev (1990) and Baillie and Bollerslev (1990), for example, use a multivariate GARCH model on weekly data, but do not achieve strong results.

The nonparametric treatment of the conditional variance $\sigma^2_{t+1|t}$ is motivated by the limitations of a linear specification for the mean equation in the ARCH-M model. A nonlinear mapping between the conditional variance and the information set is more likely to be captured in a flexible context (cf. Pagan and Hong, 1991). An explicit parameterization of the risk term introduces uncertainty about the interpretability of the results because of the possible misspecification of the model, or of some undesirable properties (such as persistence in the present context) in the estimated conditional variance.

Pagan and Hong (1991) have proposed to estimate flexible forms for the ARCH-M model in a nonparametric fashion on monthly data. In what follows we will discuss the instrumental variable procedure and suggest an alternative way to select the instrument for the risk–related term. Our suggestion differs from the estimators proposed by Pagan and Ullah (1988) and by Pagan and Hong (1991) in that we motivate our estimator on the ground of orthogonality conditions. Accordingly, $\sigma^2_{t+1|t}$ can be substituted by an estimable counterpart ϕ_{t+1} such that $E(\phi_{t+1}|\Psi_t) = \sigma^2_{t+1|t}$. In particular, some residuals $\hat{\epsilon}^2_{t+1}$ can be used as ϕ_{t+1}, without affecting the asymptotic properties of the estimator (Pagan and Ullah, 1988). By so doing, though, we will incur in the generated regressor problem (Pagan, 1984) since, by appropriate algebra, we see that

$$\Delta s_{t+1} = \beta_0 + \beta_1(f_t - s_t) + \delta\hat{\epsilon}^2_{t+1} + (\sigma^2_{t+1|t} - \hat{\epsilon}^2_{t+1})\delta + \epsilon_{t+1}, \qquad (5)$$

that is,

$$\Delta s_{t+1} = \beta_0 + \beta_1(f_t - s_t) + \delta\hat{\epsilon}^2_{t+1} + \epsilon^*_{t+1}.$$

It is clear that the OLS estimator is inconsistent, and that appropriate instruments are to be sought.

The solution for a model with a risk term requiring instrumental variable estimation was first suggested by Pagan and Ullah (1988) where a nonparametric estimate of the variance is used as an instrument for $\hat{\epsilon}^2_{t+1}$. The relevant issue becomes then one of the choice of optimal instruments. A semi–parametric instrumental variable estimator can be derived following Newey (1990) or Robinson (1991).

In a Generalized Method of Moments framework, the question can be posed in terms of deriving the appropriate conditions for the conditional first moment in our model. Given the equation for the mean this can be written as

$$E\left(u(\Delta s_{t+1}; \beta_0, \beta_1, \delta | \mathbf{\Psi}_t)\right) = 0$$

where u is to be seen here as the disturbance term from our model, upon conditioning on the relevant information set $\mathbf{\Psi}_t$. In unconditional terms this relationship postulates the existence of some function $\mathbf{w}(\mathbf{\Psi}_t)$ such that the following orthogonality conditions hold

$$E\left(u(\Delta s_{t+1}; \beta_0, \beta_1, \delta)\mathbf{w}(\mathbf{\Psi}_t)\right) = 0$$

The optimal (relative to a class imposing restrictions just on the first moment – cf. Newey 1990) instruments are chosen as

$$\mathbf{w}_t^{opt} = \frac{\hat{\mathbf{q}}_t}{\hat{\omega}_t} \quad \text{where} \quad \omega_t = E(u(\cdot)^2|\mathbf{\Psi}_t) \quad \text{and} \quad \mathbf{q}_t = E\left(\frac{\partial u_t(\cdot)}{\partial \boldsymbol{\theta}'} \mid \mathbf{\Psi}_t\right) \tag{6}$$

with $\boldsymbol{\theta}' = (\beta_0\, \beta_1\, \delta)$. The estimation strategy then would follow three–steps:

1. we first estimate the spot/forward relationship (without risk term) by OLS, obtaining the residuals on which a nonparametric estimation of the conditional variance is based (we defer to the Appendix some technical details about how this is performed);

2. the second step is a first round of GMM estimation, choosing $\omega_t = 1$ and deriving a robust variance–covariance matrix of the parameter estimates;

3. finally, we can derive an efficient GMM estimation, using estimated ω_t constructed on the basis of estimated residuals at step 2.

It is of course possible to think of imposing further restrictions (on conditional second moments, but also on third, fourth) thus enriching the set of orthogonality conditions on the basis of which the estimated $\boldsymbol{\theta}$ is derived. By verifying the analytical conditions provided by Newey (1993) it would be possible to analyze the relative gain in efficiency for the parameters of the mean equation. This is not pursued here, though.

4 The Limits of the Parametric Specification

The parametric specification of the GARCH class of models for the problem at hand is based on the information contained in the data sampled at a single frequency (e.g. weekly or monthly). Moreover, when the data are sampled at a higher frequency than the maturity of the forward contract, the error term in the relationship can be shown to follow an MA process of order equal to the number of sample periods included in the maturity, implying the need for a modification in the estimation procedure (cf. Gallo and Pacini 1995).

This may be seen as a downside of the methodology, particularly when applied to financial series where the flow of information available is continuous. Working with monthly data, as we do here, reflects a peculiar view of the mechanisms at work on the markets, whereby the most recent point of reference in an autoregressive framework is a month earlier and what occurs within the month would not affect the agents' perception of the situation and their decisions. We will return on this limitation in the next section, where we discuss the expansion of the information set to include within-month measures of volatility and trading signals from technical analysis.

The first set of results (cf. Tables 1 to 6 show the estimates with standard errors in parentheses) refers to a parametric specification for the six currencies (French Franc, Italian Lira, British Pound, Japanese Yen, Canadian Dollar, and US Dollar *vis-à-vis* the Deutsche Mark), where the simple spot/forward relationship is first estimated by OLS and a remaining structure (ARCH(1)) in the residuals is tested for by means of a Lagrange multiplier test (critical value at 5% = 3.84). The parameters are all insignificant across currencies, with the coefficient for the forward premium being different also from 1 (with the exception of the Yen). A Ljung-Box(12) statistic (critical value at 5% = 21.02) is computed to check autocorrelation in the residuals. The OLS results signal the need for a richer dynamic specification for the Franc, the Lira and the Yen.

For all the currencies is the null of no ARCH rejected, so that the specification is augmented by ARCH and GARCH structures for the error term on the one hand, and then extended to a specification where the conditional volatility term is included in the equation for the mean. The results for the (G)ARCH models show that the constant and the forward premium become significant for the French Franc only, with the coefficient for the forward premium getting closer tc one. As for the (G)ARCH-M specifications, the addition of the conditional variance term in the mean does not add significantly to the likelihood function, and in any case, the resulting coefficients are far from the $(0, 1)$ null hypothesis implied by the theory (with the exception, again, of the Yen, due to the imprecise estimation). Three parameters on the forward premium are negative (Pound, CA\$ and US\$). The estimated skewness and kurtosis (not reported) are such that normality is rejected for all currencies, even after taking ARCH effects into consideration.

Table 1: French Franc/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	-0.03	-0.35	-0.31	-0.34	-0.32
($\times 10^2$)	(0.14)	(0.08)	(0.06)	(0.08)	(0.06)
$(f_t - s_t)$	0.34	0.78	0.76	0.08	0.68
	(0.16)	(0.08)	(0.08)	(0.10)	(0.08)
Cond. Volat.				-0.08	0.41
				(0.13)	(0.45)
α_0		0.04	0.005	0.04	0.004
($\times 10^3$)		(0.004)	(0.001)	(0.004)	(0.001)
α_1		1.20	0.71	1.22	0.74
		(0.16)	(0.14)	(0.16)	(0.03)
γ_1			0.27		0.24
			(0.06)		(0.05)
Loglik	695.39	743.50	755.02	744.16	755.29
MSE ($\times 10^4$)	2.14	2.18	2.17	2.16	2.22
AR(12)	43.00	15.72	13.41	16.02	12.65
ARCH(1)	13.28	0.28	0.78	0.61	0.43

Table 2: Italian Lira/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	0.21	-0.16	0.004	-0.02	-0.10
($\times 10^2$)	(0.22)	(0.21)	(0.23)	(0.20)	(0.29)
$(f_t - s_t)$	0.26	0.39	0.34	0.52	0.29
	(0.18)	(0.22)	(0.22)	(0.20)	(0.20)
Cond. Volat.				-0.50	0.49
				(0.27)	(0.43)
α_0		0.23	0.03	0.24	0.03
($\times 10^3$)		(0.009)	(0.007)	(0.009)	(0.006)
α_1		0.57	0.80	0.45	0.80
		(0.12)	(0.06)	(0.10)	(0.03)
γ_1			0.09		0.09
			(0.02)		(0.02)
Loglik	594.30	616.32	620.98	617.48	620.54
MSE ($\times 10^4$)	4.82	4.92	4.86	5.50	4.76
AR(12)	21.65	10.93	8.14	10.16	8.77
ARCH(1)	4.85	0.21	0.92	0.25	0.47

Table 3: British Pound/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	0.59	0.45	0.40	0.15	0.32
$(\times 10^2)$	(0.34)	(0.34)	(0.40)	(0.51)	(0.60)
$(f_t - s_t)$	-0.77	-1.01	-0.76	-1.04	-0.82
	(0.61)	(0.69)	(0.80)	(0.69)	(0.80)
Cond. Volat.				0.35	0.13
				(0.46)	(0.51)
α_0		0.53	0.18	0.56	0.19
$(\times 10^3)$		(0.04)	(0.05)	(0.04)	(0.06)
α_1		0.35	0.57	0.32	0.55
		(0.11)	(0.10)	(0.10)	(0.11)
γ_1			0.20		0.20
			(0.06)		(0.07)
Loglik	520.06	529.31	527.75	529.81	528.03
MSE $(\times 10^4)$	8.82	8.90	8.86	8.74	8.83
AR(12)	17.85	18.75	18.50	18.56	18.51
ARCH(1)	4.59	0.27	3.06	0.41	3.46

Table 4: Japanese Yen/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	-0.31	-0.32	-0.32	5.31	0.25
$(\times 10^2)$	(0.20)	(0.20)	(0.19)	(13.51)	(0.85)
$(f_t - s_t)$	0.64	0.65	0.71	0.60	0.67
	(0.52)	(0.54)	(0.58)	(0.53)	(0.55)
Cond. Volat.				-4.45	-0.50
				(10.75)	(0.73)
α_0		0.96	0.67	1.01	0.06
$(\times 10^3)$		(0.09)	(0.65)	(0.10)	(0.05)
α_1		0.06	0.86	0.03	0.87
		(0.06)	(0.08)	(0.07)	(0.07)
γ_1			0.06		0.06
			(0.03)		(0.03)
Loglik	495.37	496.04	500.22	497.30	500.53
MSE $(\times 10^4)$	10.77	10.77	10.77	10.62	10.69
AR(12)	21.33	19.71	17.75	19.00	17.54
ARCH(1)	0.72	1.11	1.81	0.48	1.92

Table 5: Canadian Dollar/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	0.51	0.46	0.60	-0.63	1.35
$(\times 10^2)$	(0.32)	(0.32)	(0.32)	(1.47)	(1.36)
$(f_t - s_t)$	-0.87	-1.01	-1.17	-0.88	-1.24
	(0.76)	(0.75)	(0.76)	(0.74)	(0.75)
Cond. Volat.				0.69	-0.48
				(0.93)	(0.88)
α_0		1.05	0.42	1.09	0.400
$(\times 10^3)$		(0.14)	(0.45)	(0.14)	(0.41)
α_1		0.17	0.56	0.14	0.57
		(0.11)	(0.41)	(0.10)	(0.37)
γ_1			0.11		0.10
			(0.09)		(0.08)
Loglik	470.85	472.52	472.58	472.63	472.74
MSE $(\times 10^4)$	13.13	13.13	13.13	13.09	13.14
AR(12)	16.54	15.74	16.91	16.16	17.13
ARCH(1)	1.17	0.54	0.57	0.22	0.95

Table 6: US Dollar/DM

	OLS	ARCH(1)	GARCH(1,1)	ARCH-M(1)	GARCH-M(1,1)
Constant	0.28	0.34	0.20	-0.36	-0.12
$(\times 10^2)$	(0.25)	(0.25)	(0.25)	(1.65)	(2.46)
$(f_t - s_t)$	-0.81	-1.27	-0.73	-1.22	-0.86
	(0.63)	(0.68)	(0.64)	(0.66)	(0.66)
Cond. Volat.				0.49	0.24
				(1.13)	(1.68)
α_0		1.05	0.15	1.06	0.14
$(\times 10^3)$		(0.12)	(0.33)	(0.12)	(0.37)
α_1		0.12	0.82	0.11	0.84
		(0.11)	(0.32)	(0.10)	(0.34)
γ_1			0.03		0.02
			(0.05)		(0.04)
Loglik	476.59	477.58	477.12	477.70	477.28
MSE $(\times 10^4)$	12.53	12.54	12.52	12.51	12.51
AR(12)	15.87	16.01	14.46	16.08	14.62
ARCH(1)	1.66	2.12	2.15	1.81	1.83

The coefficients of the risk term are never significant, a result which is consistent with the outcome of the ARCH-M specification chosen in the paper by Domowitz and Hakkio (1985). In fact, their evidence is very similar to the one produced above (in their case, monthly data vis-à-vis the US$), since the parameters are characterized by an overall lack of significance for the mean equation, even after accounting for the ARCH specification for the error term and the inclusion of the conditional variance in the mean equation.

To a closer inspection, though, the degree of persistence in the estimated conditional volatility can be recognized as being very high, making the GARCH specification estimated here of the integrated type or close to it. Since the dependent variable is covariance-stationary (and a very short-memory one), this could account for the lack of significance of the risk term in the mean equation here and in Domowitz and Hakkio's case. Whether this is due to the autoregressive structure of the parametric specification, the presence of regime shifts (cf. Lamoureux and Lastrapes, 1990) or the assumption of normality for the disturbances will not be analyzed here, since we will shift the attention to the nonparametric estimation of the conditional variance which will avoid all the above problems.

5 The Evidence from the Semiparametric Specification

It is by now established in the literature (cf. for example Hamilton and Susmel 1994) that the (G)ARCH specification, in any of its many versions entails a degree of persistence in the conditional variance which is too high to be adequate, for example, to represent forecast confidence intervals for conditional volatility, since the effect of a high shock dies out too slowly, when compared with subsequent measures of historical volatility.

Contrary to a (G)ARCH-M specification, a nonparametric measure of conditional volatility allows us to exploit the local approximation properties of the kernel estimator, not forcing the evolution of the conditional volatility to follow a difference equation with estimated roots close to one.

The results obtained are presented in Table 7 where the risk term is estimated on the basis of residuals obtained from the spot/forward relationship estimated with monthly data. The number of lags included in the nonparametric regression, selected on the basis of information criteria, is also reported. Below each coefficient we report the robust standard errors computed on the basis of the estimator's variance-covariance matrix

$$\widehat{\text{asy.var}}(\hat{\theta}) = \left((\mathbf{W'X})\,(\mathbf{W'\Omega W})^{-1}\,(\mathbf{X'W})\right)^{-1}$$

where $\mathbf{W}$ is the matrix of instruments, the t-th row of which was derived in (6), and the ω_t's in (6) are the diagonal elements of the $\mathbf{\Omega}$ matrix.

On the basis of these standard errors, the only currencies which exhibit a significant risk–related effect are the Pound and the Yen. All the other coefficients are not significant. Some problems with autocorrelation and ARCH effects are still present for Franc and Lira, together with a rejection of normality (not reported) for all currencies.

Table 7: Semiparametric Specification
Instruments Estimated on Monthly Data

	FF	Lit	BP	JY	CA$	US$
Constant ($\times 10^2$)	-0.13	0.30	-1.23	0.99	-1.31	-1.46
	(0.10)	(0.44)	(0.96)	(0.67)	(3.35)	(1.77)
$(f_t - s_t)$	0.17	0.19	-1.60	0.37	-1.24	-1.18
	(0.42)	(0.35)	(1.29)	(0.33)	(1.75)	(1.35)
Cond. Volat.	0.75	0.04	2.17	-1.04	1.29	1.16
	(0.70)	(0.65)	(0.68)	(0.51)	(2.15)	(1.17)
MSE ($\times 10^4$)	1.67	4.20	22.99	12.90	24.95	22.49
AR(12)	57.8	29.5	8.3	8.9	7.5	9.2
ARCH(1)	39.54	8.21	0.15	0.004	0.18	0.27
Lags in Nonpar. Regression	1	2	2	1	2	1

The performance on the semiparametric specification adopted is slightly better than the parametric one but (with the mentioned exceptions) there is not a clear cut evidence of the relevance of a risk–related term in the spot/forward relationship. This leads us, as mentioned before, to investigate another aspect of the problem, that is the amount of information contained in a series sampled at a relatively low frequency such as the monthly one. In fact, an advantage of our semiparametric estimator is that we can adopt an instrument for $\hat{\epsilon}^2_{t+1}$ derived nonparametrically from data sampled at a different frequency from the one for the main model. The results with the instrument estimated at the higher frequency (weekly) and sampled at monthly intervals are reported in Table 8.

Table 8: Semiparametric Specification
Instruments Estimated on Weekly Data

	FF	Lit	BP	JY	CA$	US$
Constant ($\times 10^2$)	-0.12	0.006	0.04	0.37	0.54	0.22
	(0.10)	(0.39)	(0.36)	(0.28)	(0.55)	(0.41)
$(f_t - s_t)$	0.01	0.13	-1.03	0.50	-0.96	-0.87
	(0.24)	(0.43)	(0.71)	(0.32)	(0.98)	(0.86)
Cond. Volat.	1.05	0.59	0.66	-0.55	-0.01	0.003
	(0.003)	(0.28)	(0.18)	(0.25)	(0.31)	(0.30)
MSE ($\times 10^4$)	1.63	4.38	8.65	10.89	13.09	12.53
AR(12)	63.1	23.5	17.2	13.4	18.0	16.9
ARCH(1)	33.72	0.31	1.82	0.002	2.18	1.46
Lags in Nonpar. Regression	1	1	1	3	1	1

In spite of the lack of significance of the constant and the parameter on the forward premium, the picture on the risk term is much sharper now, since the parameters for four out of the six currencies are significant (CA$ and US$ being the exceptions), and of comparable size. This result is not surprising for the European currencies, since the overall historical behavior of these currencies *vis-à-vis* the Deutsche Mark followed a secular trend which strengthened the German currency over the years, contributing to its reputation as a safe haven within Europe. From these results one would tend to conclude that there is no fixed component of the risk-premium, but just the time-varying one. The lack of significance of the risk-premium for the North American currencies is somewhat puzzling since we know that there have been periods in the sample considered when they have been strongly appreciating or depreciating. The hypothesis that the estimated coefficients capture a sort of average effect is investigated in the next section.

6 Trading Signals and Risk Perception

The evaluation of the impact of the risk–premium on exchange rate movements thus far has shown that interest rate differentials seem to play little or no role in the short-run. However, the question of how risk is perceived across periods remains still open given that in the floating exchange rate experience since 1973 currencies have often alternated phases of strength and weakness which should correspond to a different risk perception about detaining a certain position. As Engel and Hamilton (1990) have shown, however, even when these long-term movements are taken into account the interest rate differential does not seem to convey a relevant message about future movements.

The mechanisms at work in the markets have received an increasing attention in the academic literature, since the lack of support for an expectation-based theory such as uncovered interest rate parity has raised the question as to whether considering heterogeneity is likely to make a difference when the hypothesis of rational behavior are investigated. In this respect, it is interesting to see whether certain "fads" which are generated on the markets reflect a consensus of opinions about the direction (appreciation or depreciation) taken by a currency in the short-run.

Technical analysis consists of several statistical techniques and rules of thumb which are widely followed by traders to determine their short-run position on the markets (Taylor and Allen 1992). The consequences of this "habit" is that the process of expectation formation relies much more on asymmetric information and the possibility of expectational errors, or of waves of herding mechanisms as the outcome of reciprocal influence by markets' participants (Lehmann 1990; Kirman 1993). Perhaps oversimplifying the situation, the trading rules (or filters) suggested by technical analysis translate into advice to buy or sell. Whether these signals are followed or not in practice depends on the complete set of information available to traders, and, ultimately, on what is defined as market sentiment, i.e. a sort of collective feeling about what is likely to happen.

In the present context, we are not pursuing a strategy of detecting possible pockets of profitability on the foreign exchange markets, but we want to investigate whether the analysis of trading signals helps us in identifying periods marked by a definite (and recognizable) tendency of the currency. Across periods with different tendencies the perception of risk relative to that currency must change and should be empirically detectable.

We look at these effects by characterizing various market situations on the basis of signals referred to by technical analysts as "buy", "sell", or "stay neutral" and focusing to the exchange rate movements according to which signal was prevalent on the market at each point in time. In what follows we have chosen two simple rules known as Moving Average (MA) and Moving Variance (MV) selecting periods of "buy" or "sell" when both rules signaled the same advice, and gathering into a hold-the-position period all the others (hence pooling neutral and mixed signals periods). Other, more complex, trading rules could be chosen but for the purpose of the present paper we prefer to show how even with a simple combination of signals the ensuing regimes selected have an economic interpretation.

In detail the two rules are built as follows:

1. The first rule gives out a "sell" signal (for the currency and a buy for the DM) when the period is characterized by a short-term moving average of s_t that is higher than the long-term one, and the other way around for a "buy" signal. Here we used the observation itself as the short-term average while the long-term is chosen here to contain 10 observations, although historical profitability may suggest the selection of an optimal length of the averages.

2. The second rule is based on short- and long-term moving variances

of the exchange rate forward returns defined in our case as

$$MVS_t = \frac{1}{10}\sum_{j=0}^{9}(s_{t-j} - f_{t-j-1})^2, \quad ,MVL_t = \frac{1}{100}\sum_{j=0}^{99}(s_{t-j} - f_{t-j-1})^2.$$

Action is called for when $MVS_t < (1+\alpha)MVL_t$, i.e., the short–term volatility is lower than the long–term one up to a proportionality factor α. The number of terms in each moving variance and the choice of α can be optimized (in ex post terms) when the profitability of the rule is being investigated. In this context our empirical results are based on a value of $\alpha = 0$, since experiments with a different threshold did not provide an appreciably different selection of periods. Again, a "sell" regime is characterized by periods when the previous return was positive (hence depreciation of the currency *vis–à–vis* the DM), and a "buy" regime by periods when the previous return was negative. Contrary to the previous one, for this trading rule there exist "neutral" periods as well where one should hold the position. For the purpose of the present analysis we computed the trading signals on the basis of weekly data, in order to characterize a "within-month" market situation and to lose less data at the beginning of the sample period.

In our analysis, we pool the signals from the two filters by assuming "sell" or "buy" periods characterized by consensus from the technical analysis instruments. The periods where the two signals disagree are labeled as "neutral". In order to give an idea about their characteristics we report in Table 9 some descriptive statistics by currency computed on each of the two types of periods, to be compared with each other and with the values of the whole sample period.

The descriptive results show that the differences in returns across regimes are considerable, being positive for sell periods (depreciation) and negative for buy periods (appreciation). The returns in the neutral period are hidden inside the overall effects which is a weighted average of the three types of periods. However, the exam of the minimum and maximum shows that the overall highest returns (in absolute value) occurred in a neutral period. This is not surprising because the moving variance trading rule suggests action in correspondence to a lower than usual level of recent unconditional variance, thus showing a preference for less volatile periods. This is also appreciable by the measure of variability of returns since the standard deviations are lower for the buy and sell periods relative to the overall values. The correlations between the forward premium and the conditional volatility term are very low giving support to the idea that the latter adds new information relative to the interest rate differential. Finally, the last rows of Table 9 show that the number of selected subperiods is fairly relevant adding up to more than a third of the total sample size, and that there is considerable movement in and out of each type of period (switching).

Table 9: Descriptive Statistics.

Exch.Rate Returns	FF	Lit	BP	JY	CA	US
Mean (S)	0.90	2.42	2.04	1.43	2.89	2.34
Mean (B)	-0.35	-0.32	-0.87	-1.99	-1.71	-2.16
Overall	0.14	0.45	0.28	-0.32	0.20	0.06
Std (S)	1.04	1.85	2.04	1.44	2.21	2.23
Std (B)	0.79	0.79	1.83	2.24	2.60	2.80
Overall	1.47	2.20	2.98	3.30	3.64	3.55
Min (S)	-0.50	-0.72	-1.47	-1.31	-2.32	-2.43
Min (B)	-3.47	-2.95	-6.76	-9.97	-10.20	-8.84
Overall	-4.26	-9.84	-9.39	-9.97	-10.85	-11.76
Max (S)	4.55	6.15	7.92	5.01	7.91	8.01
Max (B)	1.06	1.54	2.76	2.38	5.48	5.39
Overall	6.15	12.56	11.91	10.39	10.30	8.24
Correlations						
$(f_{t,k} - s_t), \hat{\sigma}_t^2$ (S)	0.19	0.32	0.31	-0.13	0.16	-0.04
$(f_{t,k} - s_t), \hat{\sigma}_t^2$ (B)	0.10	-0.06	-0.16	0.03	0.21	-0.05
Overall	0.39	0.03	0.08	-0.11	-0.09	0.08
Months in regime (S)	28	22	34	49	49	51
Months in regime (B)	69	72	58	60	48	52
Switching to (S)	21	19	25	32	33	36
Switching to (B)	45	43	38	37	31	25
Switching to (N)	60	56	54	59	57	55

All reported statistics are multiplied by 100.

The relevance of the selecting subperiods on the basis of the trading signals can be analyzed by estimating the separate effects of the three regimes (neutral -N-, buy -B-, and sell -S-) through some appropriate dummy variables. The model (4) becomes

$$\begin{aligned} s_{t+1} - s_t &= \beta_0 D_{Nt} + \beta_0^1 D_{St} + \beta_0^2 D_{Bt} \\ &\quad + \delta\sigma^2_{t+1|t} D_{Nt} + \delta^1 \sigma^2_{t+1|t} D_{St} + \delta^2 \sigma^2_{t+1|t} D_{Bt} \\ &\quad + \beta_1 (f_{t,1} - s_t) D_{Nt} + \beta_1^1 (f_{t,1} - s_t) D_{St} + \beta_1^2 (f_{t,1} - s_t) D_{Bt} + \epsilon_{t+1}. \end{aligned}$$

In order to improve the readability of the results, we report only the values of the estimated coefficients for the buy and sell regimes due to the residual nature of the neutral period. Table 10 summarizes the results obtained with our estimator when the instruments are derived on the basis of weekly data (the results obtained with the instruments estimated on monthly data are less sharp, although similar, and are available upon request). The first remark relates to the constant components of the risk–premium: for each currency there is just one coefficient which is different from zero (Lira, Pound, CA\$, US\$ for the sell regime i.e. when the currency has a recognized tendency to depreciate, and FF and Yen for the buy regime). This might be interpreted as a sort of reputation effect of the currency by which either sort of signal entails an instantaneous perception of the risk (respectively, the advantage) connected to holding the currency when it is depreciating (respectively, appreciating). The coefficients of the forward premium are seldom different from zero, the only notable exception being the Yen, for which both regimes have significant coefficients which are also not significantly different from 1.

Most importantly, for the goals of the present analysis, the coefficients on the conditional volatility are all significant and opposite in sign across regimes. The coefficients on the sell signal are very similar to each other, varying from 1.15 (US\$) and 2.26 (FF); the confidence intervals around each of them have a nonempty intersection around 1.5. As for the coefficients for the buy signal they are negative (in coincidence with expectations since a buy period should be characterized by an appreciation of the currency) and range from −1.29 (Yen) to −2.77 (Lira). In this case, though the confidence intervals overlap separately for the group of European currencies (for values of the coefficient below -2) and for the group of non-European currencies (for values of the coefficient around -1.2). The difference in the impact of conditional volatility is interpreted as evidence of asymmetry of the time-varying component of the risk–premium. In fact, a $\diamond$symbol by column means that the sell coefficient is significantly different from the buy coefficient in modulus.

Table 10: Analysis with Trading Signals

	FF	Lit	BP	JY	CA$	US$
Constant :				$\diamond$		
Sell Regime $(\times 10^2)$	0.26	0.75	0.82	0.19	1.95	0.88
	(0.18)	(0.31)	(0.27)	(0.18)	(0.23)	(0.39)
Buy Regime $(\times 10^2)$	-0.30	-0.06	0.40	-0.83	-0.53	0.30
	(0.07)	(0.13)	(0.28)	(0.20)	(1.10)	(0.82)
Forward Premium :						
Sell Regime	0.55	0.15	-0.05	0.85	-2.19	-0.35
	(0.65)	(0.52)	(0.80)	(0.32)	(0.62)	(0.68)
Buy Regime	0.48	0.18	-0.06	1.56	3.38	-0.10
	(0.16)	(0.14)	(0.66)	(0.34)	(4.70)	(1.08)
Conditional Volatility :	$\diamond$	$\diamond$	$\diamond$	$\diamond$		
Sell Regime	2.26	1.80	1.40	1.98	1.15	1.26
	(0.47)	(0.31)	(0.28)	(0.28)	(0.17)	(0.26)
Buy Regime	-2.38	-2.77	-2.14	-1.29	-2.16	-1.62
	(0.13)	(0.39)	(0.19)	(0.17)	(0.94)	(0.32)
MSE $(\times 10^4)$	1.22	3.65	6.94	8.91	10.77	9.44
AR(12)	20.72	17.92	10.68	12.17	12.03	15.55
ARCH(1)	3.70	0.28	0.77	0.008	0.09	0.96

$\diamond$Asymmetry Test: Sum of two regime coefficients significantly different from zero at 5% sig. level.

The results for the risk term obtained in Table 10 can be evaluated (Figures 2 to 7) by showing the time profile of the risk–premium relative to the "buy" and "sell" periods. The series depicted are computed on the basis of the estimated coefficients according to the following expression:

$$RP_{t,1} = \begin{cases} RP^S_{t,1} & = & (\hat{\beta}^1_0 + \hat{\delta}^1 \hat{\epsilon}^2_{t+1}) D_{St} \quad \text{if } t = \text{sell} \\ RP^B_{t,1} & = & (\hat{\beta}^2_0 + \hat{\delta}^2 \hat{\epsilon}^2_{t+1}) D_{Bt} \quad \text{if } t = \text{buy} \end{cases}$$

with $D_{St} = 1$ when t is a "sell" period and 0 otherwise and analogously for D_{Bt}. Note that on the basis of the estimated coefficients $RP^S_{t,1}$ is always greater than zero while $RP^B_{t,1}$ is always less than zero.

We have divided up the presentation of the trading signals for the currencies under investigation in two groups (European and Non European) since for the former the trading signals may shadow actions by the markets right before or after an institutional realignment. Figures 2 to 4 represent the occurrence of such trading signals for the European group, where we superimposed the first vertical bar in correspondence to the inception of the ERM mechanism and subsequent ones representing the dates of the central parity realignments (or crises). A pattern can be isolated, where following a realignment there is a tendency to have "buy" signals (to ripe the benefits of speculative attacks) and in some instances (for the French Franc and the Italian Lira) the "sell" signal occurs at the same time as the realignment. Experiments with our model considering the different periods between successive realignments did not signal any significant difference across periods in any of the groups of coefficients. However, from the graphs it is interesting to note the stabilizing effect of the ERM: the impact of the risk-premium on the currency movements is much higher for the FF and the Lira prior to the institution of the ERM, while for the Lira and the Pound the exit from the ERM has determined a considerable increase in the impact of volatility.

For the other group we report with vertical bars the periods of high appreciation and depreciation detected by Engel and Hamilton (1990). It is interesting to notice (Figure 7) how the herding behavior in favor or against the US$ occurred in the mid-1980's with the buy signals clustered between 1981 and 1985. Interestingly enough, the buy signals (of the currency, i.e. sales of DM) tend to repeat themselves for the same period both for the Yen (Figure 5) and the Canadian Dollar (Figure 6). Although, for the sake of brevity, we do not present the complete correlation table among sell and buy signals, we can report that there is a high correlation (0.67 for the sell and 0.62 for the buy, respectively) between the Canadian and the US Dollars. For the other currencies the correlations are much lower (just for the buy signal 0.34 between Yen and CA$ and 0.35 between Yen and US$). From a graphical point of view, while the occurrence of signals for the Canadian Dollar follows the US Dollar more closely, it is interesting to notice that in correspondence of the sell signals ("talk the dollar down"), there is a time (beginning of 1986) of buy signals for the Yen. We interpret the evidence (both graphical and based on simple correlations) as pointing to the presence of shifts in currency portfolio composition.

7 Conclusions

The hypothesis asserting that the forward exchange rate is an unbiased predictor of the future spot rate has seldom received empirical support, once the nonstationarity of exchange rates is adequately taken into consideration. Among the various explanations proposed in the literature for this failure, here we have adopted the time-varying risk–augmented relationship between exchange rate movements and the forward premium.

As noted, the possibility of excessive persistence in estimated conditional volatility may be one of the reasons why the ARCH-M specification used by Domowitz and Hakkio (1985) was so disappointing in proposing the risk term as a relevant variable in the relationship. In this respect a nonparametric estimation of the conditional variance allows us to exploit the local approximation properties of the kernel estimator thus providing a much less persistent estimate of the variance. At the same time, we may use data at different frequency (in our case weekly data for the monthly model) to estimate noparametrically the instruments on the conditional volatility term. One of the results of this paper is to show that the issue of timing of the information availability is a crucial one when trying to evaluate risk–related effects or the perception of risk on the markets. In fact, we obtain a sharper picture using the instruments estimated on higher frequency data.

However, the mechanisms at work on the markets are much more complicated than what is entailed by a partial analysis based on forward premium and conditional volatility alone. Opinions on where currencies are going are exchanged continuously on the markets at a much higher frequency than the one at which we examined the issue in this context. Intuitively, though, the time–varying nature of the risk–premium and its frequent change of sign when evaluated on exchange rates show that the perception of the risk attached to holding a specific currency changes whether the currency is being perceived as appreciating or depreciating. In order to derive a measure of this perception we borrowed from technical analysis two simple trading rules which were combined together to form consistent signals to buy or to sell. On this basis we were able to estimate the effects of the risk–premium keeping separate the periods of perceived appreciation (buy) from the periods of perceived depreciation (sell). Our results show that this distinction meets an empirical support, at least as far as the impact of conditional volatility is concerned. We obtain significant coefficients (and opposite in sign across regimes) for all currencies and evidence of asymmetric effects in four out of the six currencies considered. On the other hand, the uncovered interest rate parity theory is once again not supported by the data not delivering the needed values of the coefficient on the forward premium.

The relevance of the approach considered here can be pursued further. The evidence of profitability of the rules and the evidence produced here should be combined together to provide a measure of risk perception which should be used for forecasting purposes. In this respect, it would also be interesting to evaluate different suggestions for nonparametric evaluation of conditional volatility in this context (Pagan and Ullah, 1988; Pagan and Hong, 1991) since the results on estimation alone performed in a previous version of this paper do not provide a clear-cut evidence on the superiority of either approach, other than a greater stability of the coefficients in the case of our approach.

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A Nonparametric CV Estimation

Given the model $y_t = \mathbf{x}_t'\boldsymbol{\beta} + \epsilon_t$, the nonparametric estimation of the conditional variance of ϵ_t (conditioning on the information set $\boldsymbol{\Psi}_{t-1}$) can be performed either as a regression function

$$E\left((y_t - \mathbf{x}_t'\boldsymbol{\beta})^2 | \boldsymbol{\Psi}_{t-1}\right)$$

or as a functional of regression functions

$$E\left(y_t^2 | \boldsymbol{\Psi}_{t-1}\right) - \left(E(y_t | \boldsymbol{\Psi}_{t-1})\right)^2$$

which have equivalent properties.

Given a sequence of observations $(y_t, x_t), 1 \leq t \leq T$, the goal of non-parametric regression is to find $\mathbf{m}(\mathbf{x}^*) = E(\mathbf{y}|\mathbf{x} = \mathbf{x}^*)$, where $\mathbf{m}(\cdot)$ is unknown. The estimation would be

$$\hat{m}_t(\mathbf{x}^*) = \sum_{t=1}^{T} y_t \omega_t(\mathbf{x}^*),$$

which can be interpreted as a weighted local average, where the choice of the weights $\omega_t(\mathbf{x})$ depends on the selected method of local approximation.

The adopted estimation method is the Nadaraya-Watson kernel estimator with

$$\begin{aligned} \omega_t(\mathbf{x}^*, \mathbf{A}) &= \frac{\kappa_t(\mathbf{x}^*)}{\sum_{t=1}^{T} \kappa_t(\mathbf{x}^*)} \\ \kappa_t(\mathbf{x}^*) &= (T|\mathbf{A}|)^{-1} \kappa\left(\mathbf{A}^{-1}(\mathbf{x}^* - \mathbf{x}_t)\right), \end{aligned}$$

where $\kappa(\cdot)$ is a differentiable multivariate kernel function interpretable as a density function (in our case Gaussian) and $\mathbf{A}$ is the bandwidth matrix which rules the degree of smoothness of the estimator (trade–off between variability and bias). It can be determined optimally; in our case we chose the bandwidth as proportional to the sample standard deviation (proportionality factor $T^{-1/(4+p)}$, where p is the number of explanatory variables), following the heuristic rule suggested by Silverman (1986).

B Data Issues

In determining the correct day for the future spot rate predicted by the one-month forward rate at time t, measurement error is a potential source of bias for the conclusion that forward rates fail to predict future spot rates.

In this context we used both weekly and monthly data starting from 12:00 noon quotes on the London market. For the weekly data (1079 observations, from June 1973 to January 1994), we follow Bekaert and Hodrick (1991) in selecting Fridays as the day of the week for the forward buy transaction. We determine the correct spot transaction date in the next month taking into consideration the technical aspects of the contract detailed in Bekaert and Hodrick (1991). Accordingly, monthly data (248 observations) are obtained extracting the last business day of the month as the value date which determines the corresponding spot transaction day.

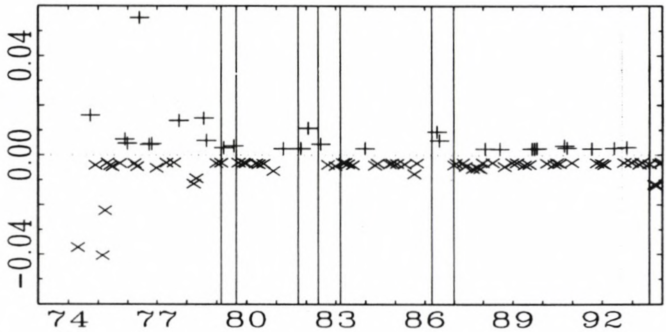

Figure 2: FF/DM - Volatility impact on exchange rate movements.

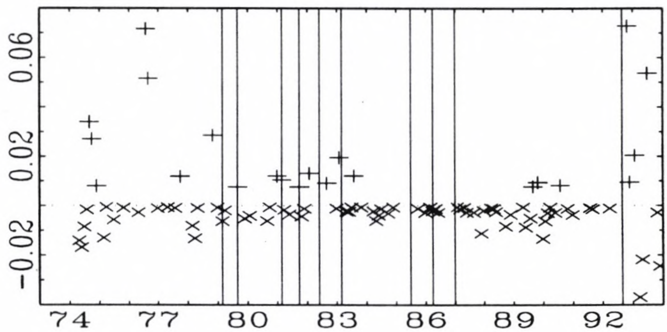

Figure 3: Lit/DM - Volatility impact on exchange rate movements.

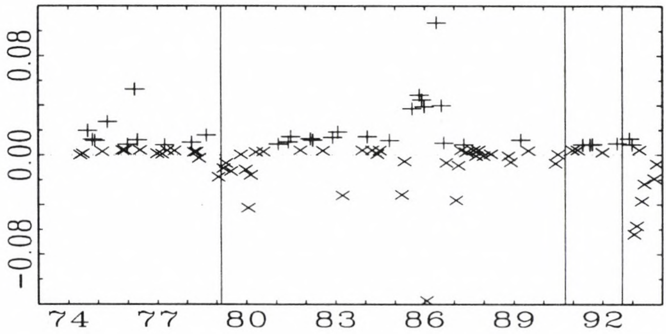

Figure 4: BP/DM - Volatility impact on exchange rate movements.

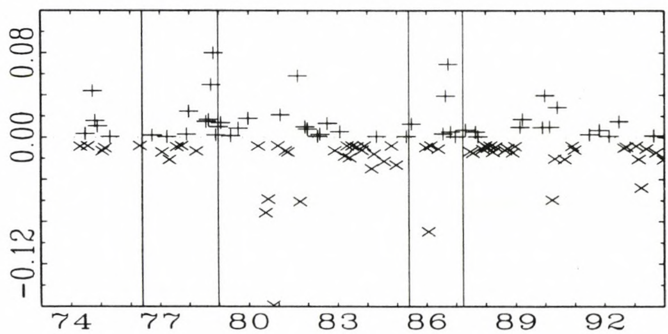

Figure 5: JY/DM - Volatility impact on exchange rate movements.

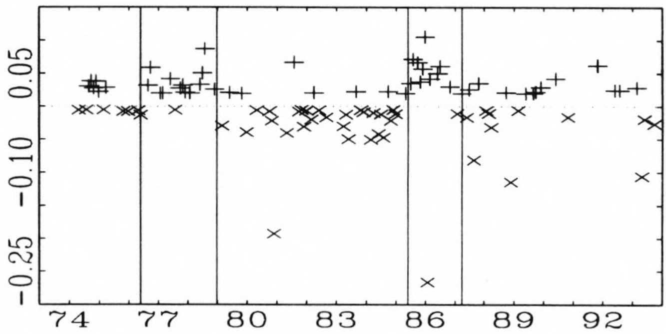

Figure 6: CA$/DM - Volatility impact on exchange rate movements.

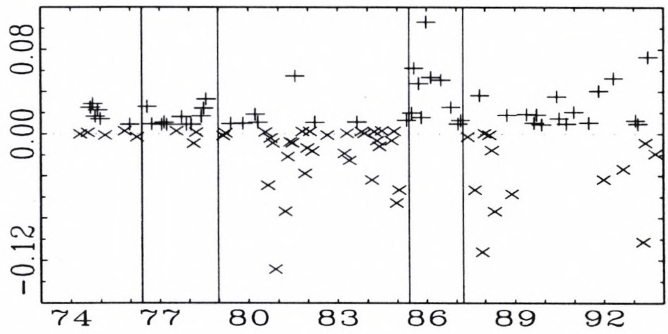

Figure 7: US$/DM - Volatility impact on exchange rate movements.

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