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The Time Variation of Expected Returns and Volatility in Foreign-Exchange Markets

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This article analyzes the time variation in conditional means and variances of monthly and quarterly excess dollar returns on Eurocurrency investments. All results are based on a vector autoregression with weekly sampled data on exchange-rate changes and forward premiums of three currencies. Both past exchange-rate changes and forward premiums predict future forward-market returns. Moreover, past forward-premium volatilities predict the volatility of exchange rates. Expected forward-market returns are very variable and persistent and exhibit marked comovements. These results carry over to cross-rate (e.g., yen/mark) investments as well.

KEY WORDS: Foreign-exchange risk premiums; Multivariate GARCH; Predictability of exchange rates; Vector autoregression.

It is well known that the forward rate is not an unbiased predictor of the future spot rate. One implication of the vast literature on the subject is that returns from investing in the forward market are predictable by the forward premium [see Hodrick (1987) for a survey and Cumby (1988) and Bekaert and Hodrick (1992, 1993) for more recent work]. Another well-known phenomenon in foreign-exchange markets is the strong volatility clustering that exchange rates display (see Bollerslev, Chou, and Kroner 1992, sec. 5, for a survey).

These empirical facts imply that conditional means and variances of forward-market returns vary through time. In this article, I document and analyze this time variation for monthly and quarterly excess dollar returns on Eurocurrency investments in the three most heavily traded currencies—the Japanese yen, the British pound, and the Deutsche mark. I use a vector autoregressive framework to address simultaneously predictability of conditional means and variances of monthly and quarterly forward-market returns using data sampled weekly. The vector autoregression (VAR) includes weekly exchange-rate changes of the three currencies and either the three-monthly or quarterly forward premiums.

The vector autoregressive approach has several advantages over the by-now-standard method of using overlapping high-frequency data on monthly or quarterly variables. First, when monthly returns are predicted using past monthly returns (or past exchange-rate changes), one implicitly weights each past weekly return with the same weight. This is not always optimal. The information present in, say, the most recent weekly return might be more useful than information in the past monthly return. For instance, when expected weekly exchange-rate changes follow a stationary (autoregressive) AR(1) process, it is easy to show that, to predict monthly exchange-rate changes, the weights on past weekly currency changes should decline. This pattern is accommodated by the VAR specification. An analogous but less general approach is that of Conrad and Kaul (1989), who used a weekly univariate model on expected stock returns to document and explain monthly expected stock returns.

Second, the use of a high-frequency model is crucial to measure the conditional variance of forward-market returns. Baillie and Bollerslev (1989) showed that the substantial serial dependence in the second moments of exchange rates, present at high frequencies, disappears at lower frequencies. Drost and Nijman (1993) ascribed this phenomenon to temporal aggregation. The weekly VAR residuals therefore allow better measurement of the true heteroscedasticity in the data. I combine the dynamics of the VAR with a multivariate model of heteroscedasticity to obtain estimates of the conditional variance of monthly and quarterly forward-market returns. In addition, I use the model to test the predictability of exchange-rate volatility by the conditional volatilities of other currencies and of forward premiums.

Third, because the forward–market returns are a function of the variables in the VAR, the dynamics of the VAR can be used to make inferences about the relationship between the returns and the information variables. This can be accomplished through the computation of implicit projection coefficients, similar to the implied VAR statistics computed by Bekaert and Hodrick (1992), Campbell and Shiller (1988), and Hodrick (1992). In addition, from the log-linear VAR specification, inferences can be made on cross-rate (e.g., yen/mark) investments as well. If similar predictability is found for cross-rates, the predictability cannot be caused by common movements of the dollar relative to other currencies, nor can it be solely due to U.S. government policies.

Fourth, the VAR contains information on expected returns on all three currency investments. In particular, conditioning on the information in the VAR, I discuss the variability and serial correlation properties of the implied expected returns. If the predictability is caused by risk, these expected returns can be interpreted as time-varying risk premiums. An interesting hypothesis in this context is the hypothesis that forward-market risk premiums are driven by a single latent variable [see Hansen and Hodrick (1983) for the original development of this test]. Although not a formal test of a fully specified model, failure to reject the latent-variable restrictions is consistent with an efficient capital market in which a limited number of common risks drive asset prices. I use the VAR framework to test these restrictions with a nonlinear Wald test. From the VAR, one can also compute correlation coefficients (and their standard errors) between the various risk premiums. As the latent-variable model in fact tests whether expected returns are perfectly correlated, correlations contain additional information about how far the data are from the null hypothesis (see also Cumby and Huizinga 1992).

The structure of the article is as follows. The first section discusses the construction of the returns and information variables, the data sources, and some summary statistics. The second section outlines the VAR framework and discusses the implied time variation in conditional means of forward-market returns. The third section focuses on the conditional variance of the returns. A final section offers some concluding remarks.

1. CONSTRUCTION AND INTERPRETATION OF VARIABLES

1.1 Definition of Returns

Consider an investment in a Eurocurrency deposit that carries an interest rate of $r^i_{t,n}$, with n the maturity in weeks and the superscript $i = 1, 2, 3$, indicating investments in yen, marks, and pounds, respectively. The holding periods considered in this article are one month (30-day contracts) and one quarter (90-day contracts). I make the simplifying assumption that a 30- (90)-day contract can be approximated by a contract over 4 (13) weeks. The analysis of Bekaert and Hodrick (1993) and the sensitivity analysis of Lewis (1990) indicated that this is a harmless assumption.

Let S^i_t be the dollar price of currency i. The uncovered dollar return on a continuously compounded Eurocurrency investment is $(S^i_{t+n}/S^i_t)\exp(r^i_{t,n})$. The rate of return in percentage points is then $100[\sum_{j=1}^n \Delta s^i_{t+j} + r^i_{t,n}]$, with Δs_t referring to *weekly* exchange-rate changes; that is, $\Delta s^i_{t+j} = \ln(S^i_{t+j}) - \ln(S^i_{t+j-1})$. The excess rate of return over a Eurodollar deposit is given by $\mathrm{fb}^i_{t+n,n} = [\sum_{j=1}^n \Delta s^i_{t+j} + r^i_{t,n} - r^4_{t,n}]$, with the number 4 indicating dollar variables. The fb symbol stands for *forward bias* because this return also corresponds to the difference between the future spot rate and the current forward rate. This follows from covered interest parity, which in continuously compounded form can be written as

$$fp^i_{t,n} = r^4_{t,n} - r^i_{t,n} \tag{1}$$

with $fp^i_{t,n} = \ln(F^i_{t,n}) - \ln(S^i_t)$, and $F^i_{t,n}$ being the forward rate for an n-week contract in dollars per foreign currency. From this it follows that $\mathrm{fb}^i_{t+n,n} = [\ln(S^i_{t+n}) - \ln(F^i_{t,n})]$, and hence $\mathrm{fb}^i_{t+n,n}$, can also be viewed as the logarithmic approximation to the return on a long forward position in the foreign currency scaled by the forward rate; that is, $\ln(S^i_{t+n}) - \ln(F^i_{t,n}) \approx (S^i_{t+n} - F^i_{t,n})/F^i_{t,n}$.

Suppressing the superscript for the foreign currency, the ex ante return in dollars to forward foreign-exchange speculation is denoted by $rp_{t,n}$:

$$rp_{t,n} = E_t\left[\sum_{j=1}^{n} \Delta s_{t+j}\right] - fp_{t,n}. \tag{2}$$

The conditional expectation will be evaluated using the dynamics of a VAR with weekly data in Section 2. Of course, there are several well-known problems with interpreting linear projections as expected returns. First, the econometrician observes only part of the information set, which gives rise to an omitted-variable problem (e.g., see Mishkin 1981; Singleton 1981). Second, the conditional expectation need not be linear. The resulting ex ante returns, however, do seem to be robust to perturbations of the information set.

The conditional second moment is defined as

$$hrp_{t,n} = E_t\left[\left(\sum_{j=1}^{n} \Delta s_{t+j} - E_t\left[\sum_{j=1}^{n} \Delta s_{t+j}\right]\right)^2\right]. \tag{3}$$

In Section 3, I employ the dynamics of the VAR coupled with a multivariate generalized autoregressive conditional heteroscedasticity (GARCH) model estimated on its residuals to compute estimates of the conditional volatility of forward-market returns.

1.2 Information Variables

The first set of information variables consists of past exchange-rate changes. As Equation (2) shows, serial correlation of exchange-rate changes implies predictability of the forward bias. Although some authors claim that exchange rates are martingales (Meese and Rogoff 1983), the success of technical trading rules (LeBaron 1991; Levich and Thomas 1993) and evidence from variance ratio tests (Liu and He 1991) indicates predictability of exchange-rate changes with respect to their own past. Table 1 also reveals some significant autocorrelations for the exchange-rate series used here.

The second set consists of the three forward premiums, which are part of the currency's risk premium. If the forward premium on a currency predicts the forward-market return on that currency and forward-market returns move together, it might also predict other forward-market returns.

To exploit all information present in monthly and quarterly interest rates, I also experimented with term structure variables (see Bekaert 1992). With the exception of the quarterly mark returns, forward premiums are by far the most important predictors of future forward-market returns. Hence I choose

Table 1. Time Series Properties of the Information Variables: Univariate Properties

	Mean	σ	ac_1	ac_2	ac_3	ac_5	ac_{13}
Δs_t^1	1.093	16.697	.094	.131	.031	.028	.052
Δs_t^2	.432	17.740	.063	.025	.031	−.061	.034
Δs_t^3	−.464	17.860	.035	.000	.049	−.040	−.007
$fp_{t,4}^1$	3.172	3.589	.951	.931	.903	.860	.691
$fp_{t,4}^2$	3.299	2.524	.969	.937	.904	.822	.561
$fp_{t,4}^3$	−2.711	3.522	.962	.936	.909	.864	.678
$fp_{t,13}^1$	3.053	3.192	.966	.951	.934	.899	.744
$fp_{t,13}^2$	3.201	2.358	.964	.938	.910	.849	.622
$fp_{t,13}^3$	−2.591	3.248	.973	.953	.931	.893	.729

NOTE: The sample period is January 1, 1975, to July 19, 1991, for a total of 863 weekly observations. Weekly logarithmic exchange-rate changes are denoted by Δs_t^i, with $i = 1, 2, 3$ indicating, respectively, the yen, mark, and pound. Similarly, $fp_{t,n}^i$ stands for the forward premium over n weeks on currency i in terms of the dollar. The number 4 indicates the dollar. All variables are annualized, forward premiums are multiplied by 1,200 (400), and weekly currency depreciation is also multiplied with 1,200. The symbol σ always denotes the standard deviation, whereas ac_n denotes the nth autocorrelation. The standard error for the autocorrelations under the white noise null is .034.

to exclude term spreads from the information set to measure conditional expected returns and conditional variances.

1.3 The Data

The data set consists of weekly observations on spot and one- and three-month forward rates of the Japanese yen, the Deutsche mark, and the British pound versus the dollar for the period of January 1, 1975, to July 19, 1991. The data are from Citicorp Data Services. All rates are sampled each Friday and are averages of bid and ask rates. When Friday was a holiday, the Thursday rate was picked. Bekaert and Hodrick (1993) explicitly examined the impact of transaction costs and the correct alignment of future spot rates with current forward rates on the measurement of foreign-exchange risk premiums. They found the bias induced by the measurement error to be very small.

1.4 Time Series Properties of the Information Variables

Means, standard deviations, and autocorrelations of the information variables are reported in Table 1. Weekly exchange rates are quite variable and show only a few autocorrelations that are significantly different from 0. All other information variables are highly autocorrelated, but the forward premiums are especially persistent. The positive (negative) forward premium means indicate that on average Japanese and German (U.K.) interest rates have been lower (higher) than U.S. interest rates.

I also report a correlation matrix for the nine variables in Table 2. The one-month and three-month forward premiums are very highly correlated, and the forward premiums on different currencies also show significant positive correlation.

2. CONDITIONAL MEANS OF FORWARD–MARKET RETURNS

2.1 A VAR Framework

Let $Y_t = [\Delta s_t^1, \Delta s_t^2, \Delta s_t^3, fp_{t,n}^1, fp_{t,n}^2, fp_{t,n}^3]'$ summarize the variables included in the VAR for $n = 4, 13$. The first important issue to be addressed is the determination of the appropriate VAR order. One potential disadvantage of this approach is that it might necessitate a large-order VAR to capture all of the correlation patterns in the weekly data. The Schwarz criterion, reported in Table 3 for the one-month holding-period specification, however, always selects the first-order VAR. The Akaike criterion, on the other hand, selects a second-order VAR, and likelihood ratio tests reject the restrictions when testing a first-order VAR versus a second-order VAR. To further investigate this issue, I report the Cumby and Huizinga (1993) l test for the presence of remaining serial correlation in the residuals of a first-order VAR (see Table 4). If the right VAR order is selected, the residuals should be serially uncorrelated. The l test is robust to both conditional heteroscedasticity and the fact that the residuals are estimated. Based on this test, the first-order VAR successfully eliminates all serial correlation in the residuals of the last

Table 2. Time Series Properties of the Information Variables: Correlation Matrix

	Δs_t^2	Δs_t^3	$fp_{t,4}^1$	$fp_{t,4}^2$	$fp_{t,4}^3$	$fp_{t,13}^1$	$fp_{t,13}^2$	$fp_{t,13}^3$
Δs_t^1	.634	.484	−.101	−.063	−.063	−.111	−.071	−.062
Δs_t^2		.708	−.072	−.080	−.120	−.079	−.079	−.116
Δs_t^3			−.039	−.031	−.090	−.055	−.039	−.084
$fp_{t,4}^1$				.774	.737	.981	.739	.726
$fp_{t,4}^2$					.718	.776	.974	.711
$fp_{t,4}^3$						.750	.698	.987
$fp_{t,13}^1$							.762	.751
$fp_{t,13}^2$								.705
$fp_{t,13}^3$								

NOTE: See note to Table 1.

Table 3. First-Order Vector Autoregression on Weekly Exchange-Rate Changes, and the One-Month Forward Premiums of the Yen, Mark, and Pound: Selection Criteria

VAR order	Akaike criterion	Schwarz criterion
1	14.608	14.536
2	14.439	14.836
3	14.468	15.066
	Likelihood ratio tests	
	1 vs. 2	214.014 (.000)
	2 vs. 3	45.884 (.125)

NOTE: The appropriate lag length for the VAR minimizes the Akaike or Schwarz criterion. The likelihood ratio test is a sequential test of a VAR(n) versus a VAR($n+1$). The test statistic has a χ^2-distribution with degrees of freedom equal to the number of coefficients being restricted by the lower VAR order. The statistic incorporates the Sims (1980) correction.

five equations, but there is still some residual correlation left in the residuals of the yen-exchange-rate-changes equation. Because estimation of an additional VAR order eliminates all serial correlation in the residuals, the remainder of the analysis is based on both first-order and second-order VAR's. To conserve space, I only report the parameter estimates for the first-order VAR.

Table 5 contains the parameter estimates for the one-month forward-premium specification for a first-order VAR. The Wald test of no predictability for weekly exchange-rate changes rejects at the 5% level for all three currencies with the predictability being the strongest for the pound. The linear predictability seems to be due primarily to the own-forward premium for the yen and pound equations, whereas in the mark equation the pound forward premium and weekly mark and pound exchange-rate changes enter significantly. Pound exchange-rate changes also help predict the future pound forward premium, and the pound and mark forward premiums help predict the future yen forward premium. In a second-order VAR (not reported), past exchange-rate changes contain information on future yen (Δs^1_{t-1}) and mark (Δs^2_{t-1}, Δs^3_{t-1}) exchange-rate changes. Although the coefficients on forward premiums are typically not significant, the predictability of yen and mark exchange-rate changes becomes stronger, both in terms of higher adjusted R^2's and lower p values for the predictability tests.

The results for the three-month-forward premium are qualitatively similar and are not reported to conserve space. The main difference is that exchange-rate changes are somewhat more important as predictors of future exchange-rate changes.

2.2 Implied Statistics for Forward–Market Returns

It is straightforward to conduct inference about forward-market returns based on the weekly VAR. To do so, all that is needed is the matrix of parameters, A, and the innovation matrix, V, of the VAR. If the VAR is of higher order, it can be written in first-order companion form, and A and V then represent the appropriately transformed parameter and innovation matrices. Let X_t be the vector of VAR variables. The first p variables in X_t are $[\Delta s^1_t, s^1_{t-1}, \ldots, \Delta s^1_{t-p+1}]'$ with p the lag order of the VAR. The other variables enter analogously, and X_t contains a total of $6p$ variables.

The first set of statistics is the coefficients of the projection of a forward-market return on the various information variables. I compute these statistics as the unconditional covariance between the forward-market return and the information variable implied by the VAR divided by the variance of the information variable. Define the unconditional variance of the VAR variables X_t to be $C(0)$. Consider the example of yen forward-market returns. Let $e1$ and $e4$ be indicator vectors of length $6p$ with zeros everywhere except for a 1 at the 1st and $(3p+1)$th spot, respectively. Thus, $e1$ selects yen depreciation and $e4$ selects the yen forward premium at time t. The vector ej analogously selects the information variable on which the yen forward bias is projected. The denominator of the projection coefficient is immediately given

Table 4. First-Order Vector Autoregression on Weekly Exchange-Rate Changes, and the One-Month Forward Premiums of the Yen, Mark, and Pound: Residual Diagnostics

Eq.	I(4)	I(13)	Q2(4)	Q2(13)	ARCH(4)	ARCH(13)	Ku	Sk	BJ
1	10.803	25.008	8.749	14.606	7.911	12.898	4.811	.664	894.73
	(.029)	(.023)	(.068)	(.333)	(.095)	(.456)	(.000)	(.000)	(.000)
2	4.775	19.353	50.300	70.838	38.296	50.923	2.374	.298	215.13
	(.311)	(.113)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
3	3.623	9.516	95.758	145.60	64.756	112.78	2.653	.044	252.99
	(.459)	(.733)	(.000)	(.000)	(.000)	(.000)	(.000)	(.598)	(.000)
4	9.633	17.202	199.43	249.40	239.19	217.38	18.10	−.032	11,761.3
	(.047)	(.190)	(.000)	(.000)	(.000)	(.000)	(.000)	(.704)	(.000)
5	4.591	11.293	306.22	776.92	173.02	310.81	10.18	−.280	3,732.7
	(.332)	(.586)	(.000)	(.000)	(.000)	(.000)	(.000)	(.001)	(.000)
6	2.418	8.494	164.89	191.56	172.04	188.93	19.96	−.577	14,357
	(.659)	(.810)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)

NOTE: The Cumby–Huizinga (1993) I test for serial correlation of the residuals is robust to conditional heteroscedasticity and lagged dependent variables. The $Q2$ test is the Ljung–Box test statistic applied to squared residuals. The ARCH test is the standard Lagrange multiplier test for serial correlation in the squared residuals, as proposed by Engle (1982). All tests are $\chi^2(n)$ with n the number of lagged (autocorrelations of) squared residuals included in the test. Ku is the normalized kurtosis coefficient, and Sk the normalized skewness coefficient. Their asymptotic distribution under the null of normality is $N(0, 24/T)$ and $N(0, 6/T)$ respectively, with T the sample size. BJ is the Bera–Jarque (1982) test for normality and is $\chi^2(2)$. P values, based on the χ^2 distribution, are reported for all test statistics.

Table 5. First-Order Vector Autoregression on Weekly Exchange-Rate Changes, and the One-Month Forward Premiums of the Yen, Mark, and Pound: Parameter Coefficients

	Const.	Δs^1_{t-1}	Δs^2_{t-1}	Δs^3_{t-1}	$fp^1_{t-1,4}$	$fp^2_{t-1,4}$	$fp^3_{t-1,4}$	R^2	$\chi^2(6)$
Δs^1_t	1.653	.08510	−.02141	.03058	−.426*	.184	−.044	.009	13.181
	(1.639)	(.05393)	(.04964)	(.04116)	(.259)	(.381)	(.220)		(.040)
Δs^2_t	−1.831	−.00305	.11646**	−.08832*	.251	−.070	−.590**	.009	15.582
	(1.834)	(.04819)	(.05938)	(.05265)	(.275)	(.448)	(.235)		(.016)
Δs^3_t	−5.960**	.05082	.03330	−.02451	.372	.466	−.991**	.013	17.973
	(1.821)	(.04786)	(.06203)	(.05817)	(.266)	(.410)	(.263)		(.006)
$fp^1_{t,4}$	.310**	.00041	−.00016	.00047	.875**	.071**	.055**	.910	5917.1
	(.129)	(.00272)	(.00322)	(.00238)	(.033)	(.034)	(.019)		(.000)
$fp^2_{t,4}$	.134*	.00144	−.00281	−.00051	.010	.956**	.009	.947	6036.6
	(.073)	(.00197)	(.00292)	(.00185)	(.011)	(.020)	(.010)		(.000)
$fp^3_{t,4}$	−.358	.00131	.00152	−.00471*	.022	.029	.930**	.927	6728.6
	(.226)	(.00273)	(.00309)	(.00273)	(.023)	(.033)	(.037)		(.000)

NOTE: Parameter estimates are obtained by ordinary least squares and reported here with heteroscedasticity-consistent standard errors. The $\chi^2(6)$ statistic tests the joint hypothesis that the six lagged variables have no predictive power. Asymptotic p values are reported in parentheses.

by $ej'C(0)ej$. The covariance term can be written as

$$\text{cov}\left(\text{fb}^1_{t+n,n}, ej'X_t\right) = e1'\,AQC(0)ej - e4'C(0)ej. \quad (4)$$

The matrix Q arises from the use of weekly data to compute monthly or quarterly depreciation [see Eq. (2)]:

$$Q = A^{n-1} + A^{n-2} + \cdots + A + I, \quad (5)$$

where I is the identity matrix. Analogously, projection coefficients can be computed for mark and pound forward-market returns.

In Table 6, I report these projection coefficients computed from a second-order VAR. The standard errors for all the coefficients reported are found using Hansen's (1982) generalized method of moments (GMM) coupled with the delta method; see Bekaert and Hodrick (1992). The projection of the monthly forward biases onto their own monthly forward premium yields coefficients varying between −2.620 (yen) and −2.825 (pound) that are significantly different from 0 except for the mark. The projection coefficients onto forward premiums of other currencies are also negative but typically carry larger standard errors with again the exception being the mark, for which the projection onto the pound premium is significantly different from 0. The projection coefficients for quarterly returns are generally somewhat larger in absolute value, but they also are somewhat more imprecisely estimated.

Past monthly or quarterly exchange-rate changes constitute the other implicit instrument in the VAR. I focus on the projection of the forward bias $\text{fb}^i_{t+n,n}$ onto $\sum_{j=0}^{n-1}\Delta s^i_{t-j}$. The numerator of this coefficient has n terms, each given by, for instance, for the yen,

$$\text{cov}\left(\text{fb}^i_{t+n,n}, \Delta s^i_{t-j}\right) = e1'A^{j+1}QC(0)e1 - e4'A^jC(0)e1, \quad j = 0, 1, 2, \ldots, n-1, \quad (6)$$

with $A^0 = I$, and the denominator is

$$\text{var}\left(\sum_{j=0}^{n-1}\Delta s^1_{t-j}\right) = e_1'[nC(0) + (n-1)(C(1) + C(1)') + \cdots + C(n-1) + C(n-1)']e_1, \quad (7)$$

Table 6. Implied Projection Coefficients of $fb^i_{t+n,n}$ Onto Variables in the Time t Information Set

Forward bias	$fp^1_{t,n}$	$fp^2_{t,n}$	$fp^3_{t,n}$	$fb^i_{t,n}$	Δs^i_t	Δs^i_{t-1}	Δs^i_{t-2}	Δs^i_{t-n}	R^2
$\text{fb}^1_{t+4,4}$	−2.620	−2.311	−1.855	.151	.0588	.0425	.0188	.0147	.083
	(.678)	(1.124)	(.767)	(.039)	(.0119)	(.0104)	(.0076)	(.0072)	(.038)
$\text{fb}^2_{t+4,4}$	−1.627	−2.624	−2.276	.082	.0291	.0178	.0137	.0117	.058
	(.859)	(1.407)	(.724)	(.038)	(.0121)	(.0115)	(.0083)	(.0080)	(.036)
$\text{fb}^3_{t+4,4}$	−1.305	−1.417	−2.825	.086	.0196	.0149	.0168	.0153	.094
	(1.031)	(1.416)	(.977)	(.055)	(.0148)	(.0132)	(.0128)	(.0122)	(.066)
$\text{fb}^1_{t+13,13}$	−2.905	−2.334	−1.957	.156	.0219	.0176	.0110	.0070	.146
	(.795)	(1.573)	(.948)	(.078)	(.0055)	(.0057)	(.0059)	(.0053)	(.081)
$\text{fb}^2_{t+13,13}$	−1.739	−2.361	−2.313	.126	.0138	.0106	.0094	.0068	.102
	(1.058)	(1.636)	(.823)	(.091)	(.0070)	(.0072)	(.0072)	(.0062)	(.087)
$\text{fb}^3_{t+13,13}$	−1.576	−1.480	−2.872	.143	.0119	.0102	.0103	.0072	.184
	(1.228)	(1.891)	(1.014)	(.127)	(.0102)	(.0099)	(.0103)	(.0081)	(.142)

NOTE: The projection coefficients of the monthly forward bias $fb^i_{t+n,n}$ ($i = 1,2,3$) onto $fp^j_{t,n}$, ($j = 1,2,3$), $fb^i_{t,n}$, Δs_{t-i} ($i = 0,1,2,\ldots,n-1$) are discussed in the text [Eqs. (4)–(7)]. An implied R^2 is also reported. The standard errors are computed as by Bekaert and Hodrick (1992). The coefficients are based on a second-order VAR for both monthly and quarterly returns.

Table 7. Properties of the Implied Risk Premiums: Standard Deviations Autocorrelations of Implied Risk Premiums

	$rp^1_{t,4}$	$rp^2_{t,4}$	$rp^3_{t,4}$	$rp^1_{t,13}$	$rp^2_{t,13}$	$rp^3_{t,13}$
σ	11.125	9.199	11.513	9.923	7.788	10.350
	(2.701)	(3.007)	(4.223)	(3.132)	(3.794)	(4.764)
ac_1	.887	.872	.935	.959	.959	.965
	(.062)	(.090)	(.035)	(.019)	(.030)	(.017)
ac_2	.786	.775	.891	.927	.926	.936
	(.110)	(.151)	(.052)	(.034)	(.052)	(.030)
ac_3	.743	.740	.851	.902	.904	.908
	(.125)	(.165)	(.065)	(.044)	(.063)	(.042)
ac_5	.681	.691	.782	.859	.865	.855
	(.139)	(.177)	(.084)	(.060)	(.081)	(.063)
ac_{13}	.528	.557	.567	.720	.738	.679
	(.153)	(.194)	(.133)	(.110)	(.134)	(.125)

NOTE: The standard deviation (σ) and autocorrelations (ac_n) of the risk premium are computed from the dynamics of the VAR. All statistics are computed from a second-order VAR.

with $C(k)$ the kth-order autocovariance of X_t, $C(k) = A^kC(0)$. Evaluating the n terms separately reveals whether the VAR weighs the different weekly components of past monthly or quarterly currency depreciation differently in predicting the forward-market return. For the yen and the mark, the weights decline with the horizon, and the most recent weekly exchange-rate change contains most information on future forward-market returns. Although the weights differ for the pound, the differences are minimal and the coefficients are imprecisely estimated. The results are similar for the quarterly forward-market returns.

Another projection coefficient I report is the projection of the forward-market return $\mathrm{fb}^i_{t+n,n}$ onto $\mathrm{fb}^i_{t,n}$. The projections on the lagged forward bias indicate that high past returns predict high future returns. The implied autocorrelations are higher for quarterly returns, and yen forward-market returns show most serial correlation.

The implied R^2 statistic reported in the table is a measure of how much of the forward bias variance is implicitly explained by the VAR variables. This can be computed as 1 minus the forecast-error variance (or equivalently the variance of the expected return) divided by the total-return variance (see also Bekaert and Hodrick 1992; Hodrick 1992). The formulas for both the projection coefficients of the forward bias on its own past and the implied R^2 are straightforward but tedious to derive and are omitted to conserve space. The VAR explains about 6%, 8%, and 9.5% of the total return variance of the mark, yen, and pound returns, respectively. The explained variance is substantially higher for quarterly returns and exceeds 18% for the pound. This is similar to the result of Hodrick (1992), who demonstrated that a relatively large amount of long-run predictability of stock returns is consistent with only a small amount of short-run predictability.

The log-linear VAR specification can also be used to make inferences on cross-rates, which have not received a lot of attention in the literature. Nevertheless, given the turbulence in fiscal and monetary policies in the United States during the 1980s, it could be postulated that the striking negative projection coefficients of forward-market returns on their own forward premiums are a "dollar phenomenon." To address this, I compute the covariance of $\mathrm{fb}^i_{t+n,n} - \mathrm{fb}^j_{t+n,n}$ with $fp^i_t - fp^j_t$ divided by the variance of $fp^i_t - fp^j_t$ for the pairs $(i, j) = (2, 1), (3, 1)$, and $(3, 2)$, representing, respectively, yen/mark, yen/pound, and mark/pound rates. The coefficients are—with standard errors in parentheses—-2.993 (.453), -4.439 (.776), and -2.495 (.987). All of these projection coefficients are significantly negative, and the negative relationship between yen/pound returns and the yen/pound forward premium is stronger than any of the projection coefficients involving the dollar.

2.3 Expected Returns

In Tables 7, 8 and 9, I summarize some properties of conditionally expected forward-market returns. The standard deviation and autocorrelations of the risk premiums are functions of the VAR parameters and are derived in the same way as the coefficients in Table 6. The standard deviation of the risk premium on both monthly and quarterly forward-market investments varies between about 8% and 11.5%. Risk premiums are not only volatile, they are also very persistent, and the autocorrelations for the monthly (three-monthly) forward-market expected return decay slowly from about .9 (.95) at

Table 8. Properties of the Implied Risk Premiums: Standard Deviations of Cross-rate Risk Premiums

	$rp^2_{t,4} - rp^1_{t,4}$	$rp^3_{t,4} - rp^1_{t,4}$	$rp^3_{t,4} - rp^2_{t,4}$	$rp^2_{t,13} - rp^1_{t,13}$	$rp^3_{t,13} - rp^1_{t,13}$	$rp^3_{t,13} - rp^2_{t,13}$
σ	9.773	12.133	7.575	8.138	10.663	5.423
	(2.392)	(2.738)	(2.143)	(2.555)	(2.999)	(2.432)

NOTE: The dynamics of the VAR are used to compute the standard deviations of the cross-rate risk premiums. All statistics are computed from a second-order VAR.

Table 9. Properties of the Implied Risk Premiums: The Comovement Between Risk Premiums

	$rp^2_{t,4}$	$rp^3_{t,4}$	$rp^2_{t,13}$	$rp^3_{t,13}$
$rp^1_{t,n}$	.551	.426	.601	.447
	(.265)	(.270)	(.303)	(.325)
$\chi^2(6)$	2.248	4.731	1.164	4.150
	(.896)	(.579)	(.979)	(.656)
$\chi^2(12)$	6.851	6.362	7.693	7.441
	(.867)	(.897)	(.809)	(.827)
$rp^2_{t,n}$		.754		.858
		(.189)		(.173)
$\chi^2(6)$		1.546		(1.436)
		(.956)		(.964)
$\chi^2(12)$		2.971		1.951
		(.996)		(.999)

NOTE: The dynamics of the VAR are used to compute the correlations between risk premiums. All statistics are computed from a second-order VAR. The $\chi^2(n)$ statistics refer to the nonlinear Wald tests of the single latent-variable restrictions. *P* values are reported in parentheses. When $n = 6$ (12), the test is from a first- (second)-order VAR. Although sensitive to the normalization of the projection coefficient ratios, the test based on the inverse of the ratio leads to the same inference.

the first lag to about .55 (.70) at the 13th lag. The standard deviation of risk premiums on cross-rate investments is also reported. The mark/pound risk premium is markedly less volatile than the risk premiums versus the dollar, but the monthly yen/pound risk premium is the most variable of all.

The time path of conditionally expected monthly returns implied by the second-order VAR is graphed in Figures 1 (dollar returns) and 2 (cross-rates). The evidence for quarterly returns and returns based on first-order VAR's looks very similar. The graphs reveal that the unconditional mean of the forward bias is close to 0 and in fact can be shown to be insignificantly different from 0. This has led some authors to argue that unbiasedness holds "on average" and that expected return differentials between covered and uncovered foreign investments are negligible (e.g., see Eun and Resnick 1988). The conditionally expected returns, however, show considerable time variation, frequently change sign, and seem very variable. The volatility of the ex ante returns seems to be lower in the eighties than in the seventies.

Figure 1 shows that the expected dollar returns from investing in the three forward markets became markedly negative during the dollar upswing in the first half of the eighties. During that period, U.S. interest rates were relatively high, although the dollar appreciated. For the mark and pound, this pattern reversed in 1985 when the dollar started to slide. Despite a positive interest differential on the dollar, expected returns from investing in the yen forward market turned positive during 1983 as well. This can be understood by turning back to Table 5, which contains the VAR results. One difference between the regression results for the yen versus the pound and mark is that the constant in the yen regression is positive. This implies that, when the interest differential is small enough, positive risk premiums on yen investments can still occur.

Figure 2 shows ex ante returns for cross-rates. The risk premiums are of the same order of magnitude and show similar patterns to the ones versus the dollar. Expected excess returns for European investors from investing in the yen Eurocurrency market were predominantly positive from 1982 till 1988. This implies for instance that Japanese stocks must have looked particularly attractive for European investors because a positive expected excess return in the Japanese stock market (see Bekaert and Hodrick 1992; Campbell and Hamao 1992) was coupled with a positive expected foreign-exchange return. For most of the floating-exchange-rate

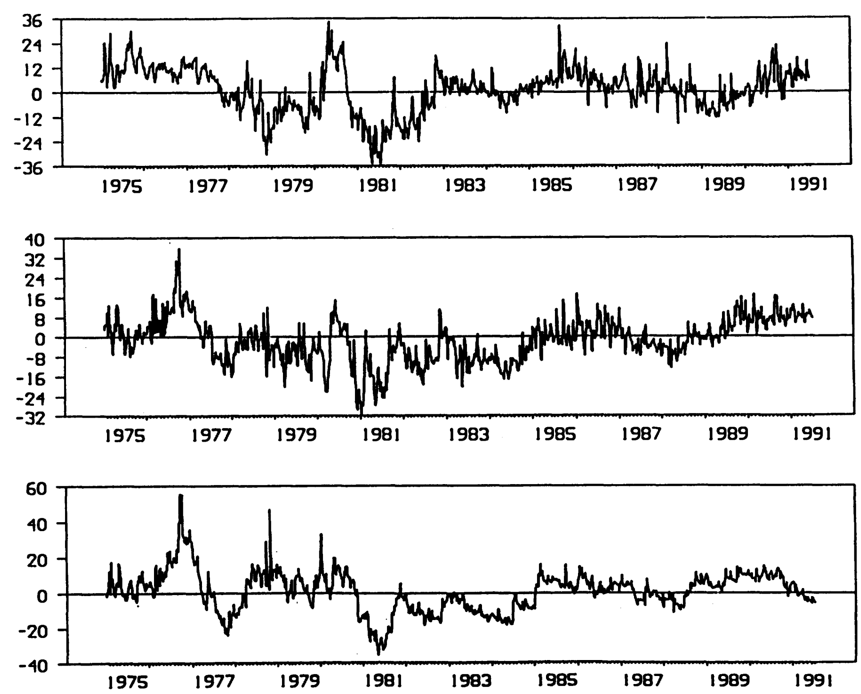

Figure 1. Ex Ante Returns Yen (data), Ex Ante Returns Mark (data), Ex Ante Returns Pound (data).

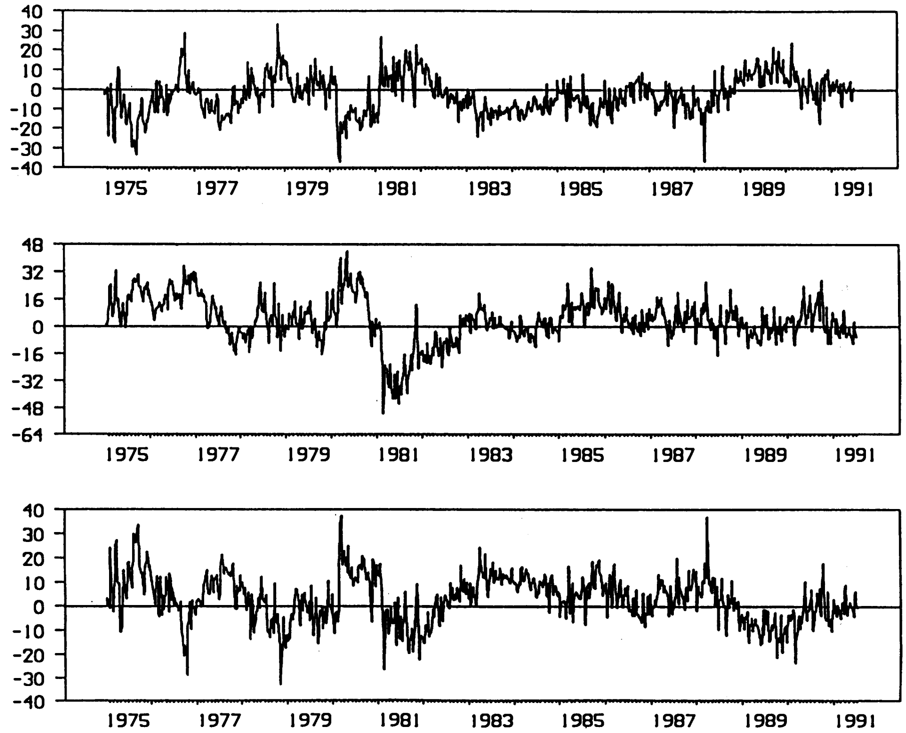

Figure 2. Ex Ante Returns Yen/Mark (data), Ex Ante Returns Yen/Pound (data), Ex Ante Returns Pound/Mark (data).

period, expected returns from investing in the pound forward market were positive for mark-based investors because the pound–mark interest differential has generally been positive.

2.4 Comovement of Expected Returns

A final important question is whether the risk premiums on different currencies move together. Figure 1 clearly shows similar time variation in all three expected returns, although the comovement between the yen and the other currencies seems to be less strong. The single latent-variable model essentially predicts perfect correlation between expected returns. Therefore, the null hypothesis can be written as

$$rp_t^i = \beta rp_t^j, \qquad i \neq j. \tag{8}$$

Consider projecting two different returns on a set of instrumental variables Z_t, and denote the projection coefficients by (α_1, α_2). The latent-variable model then imposes the restriction $\alpha_1 = \beta\alpha_2$. The standard latent-variable test is carried out by estimating (α_1, β) with the GMM and using the standard test of the overidentifying restrictions to test the null. Lewis (1990) showed that, as the maturity horizon increases, the model tends to be rejected less often. Typically the restrictions are not rejected for three-month forward-market returns, but the test rejects for monthly returns. In the context of this article, I attempt to infer information on the null of Equation (8) for both monthly and quarterly forward-market returns using the weekly VAR specifications.

The first test I consider is a nonlinear Wald test of the null hypothesis. The VAR implicitly contains $6p + 1$ information variables (with p the VAR order), $Z_t = [1, X_t']'$. The VAR can be used to compute the implicit regression coefficients of the various forward-market returns onto Z_t. Let's denote the vector of regression coefficients for $\mathrm{fb}_{t+n,n}^i$ by α_i. Under the null of the single latent-variable model, all of the elements of the vector containing the ratios of the elements in α_i and $\alpha_j (i \neq j)$ should equal β. This imposes $6p$ restrictions, which can be tested with a nonlinear Wald test, derived in an appendix available from the author.

The test has two disadvantages. One is that it does not recover an estimate for the factor of proportionality β as the GMM test does. Second, the test is sensitive to the way the ratio of projection coefficients is taken because this involves a nonlinear transformation. Fortunately, the inference does not change when the ratio is reversed. To its advantage, the test avoids the nonlinear GMM estimation, which often has dubious small-sample properties (see Ferson and Foerster 1994). The test results are given in Table 9. The test does not reject for any of the pairs of expected monthly and quarterly returns.

As Cumby and Huizinga (1992) stressed, computing the correlation coefficient between two expected returns provides additional insights in the match between data and the null hypothesis of Equation (8). From Table 9, the correlations between yen expected returns and both pound and mark expected returns appear less than perfect in an economic sense, but they are imprecisely measured. The correlation between mark and pound risk premiums is close to 1 in both the statistical and economic sense. This raises the question of whether similar forces drive European currencies that are not present for the yen (and vice versa). Because the pound only entered the European Monetary System at the very end of the sample period, a different exchange-rate regime is not behind the results. It is also striking that, contrary to Lewis's (1990) results, the evidence for monthly and quarterly returns is so similar.

Table 10. Parameter Estimates From a Multivariate GARCH Model: Likelihood Ratio Tests

DIAG (60)	CM (48)	FPDS I (36)	FPDS II (18)	FPDS III (18)
312.8	275.4	150.2	120.1	75.6
(.000)	(.000)	(.000)	(.000)	(.000)

NOTE: Parameters of the BEKK model in Equation (10) in the text are obtained by maximum likelihood using a multivariate normal distribution for the innovations of the second-order VAR on monthly forward-market returns. Under very weak conditions including misspecification of the distribution function (see Bollerslev and Woolridge 1992; White 1982), the vector of parameters is asymptotically normally distributed with covariance matrix $\Omega = A^{-1}BA^{-1}$, where A is the Hessian form and B the outer product form of the information matrix. Various likelihood ratio tests are reported. DIAG is a diagonal model, restricting each conditional variance to depend only on the own past conditional variance and squared innovation. CM also introduces the conditional variance and squared innovation of the own forward premium (exchange-rate change) in each exchange-rate (forward-premium) volatility equation. FPDS II (FPDS III) imposes no influences of conditional (co) variances of forward premiums (exchange rates) in exchange-rate (forward-premium) equations. FPDS I imposes both of these restrictions. The test statistics are all $\chi^2(n)$, where n is the number of restrictions indicated in parentheses below the model name.

3. CONDITIONAL VARIANCES OF FORWARD–MARKET RETURNS

3.1 Heteroscedasticity and Normality Tests

To test for the presence of heteroscedasticity in the VAR residuals, I perform two tests. The $Q2$ test is the Ljung–Box test applied to squared residuals, which should be serially uncorrelated under the null of conditional homoscedasticity. The autoregressive conditional heteroscedasticity (ARCH) test is the standard Lagrange multiplier (LM) test for ARCH effects of order p, found by multiplying the sample size T with the R^2 of the regression of the squared residuals on p of its lags plus a constant (see Engle 1982). The tests are χ^2 of order p, where p is the number of included (autocorrelations of) residuals. They are contained in Table 4. The test is not robust to conditional leptokurtosis (see Bollerslev and Woolridge 1992). Given the strong nature of the results, it is unlikely that a robust test would yield different inference. The reported kurtosis and skewness coefficients are both 0 under the null. The BJ statistic is the Bera–Jarque (1982) test for normality and has a $\chi^2(2)$ distribution. It is found by adding up the squared standardized kurtosis and skewness coefficients. All test values have their asymptotic p values reported underneath.

Normality is rejected for all equations. Usually this is due to excess kurtosis, but the yen and mark exchange-rate changes show positive skewness, whereas the mark and pound forward-premium residuals display negative skewness. Both negative and positive skewness make it difficult to capture all of the time variation of second moments with a pure GARCH model with normal innovations. The fat tails might well arise from heteroscedasticity because the null of conditional homoscedasticity is firmly rejected, not only for exchange-rate changes but also for the forward premiums. The only exception is the yen exchange-rate equation, in which there is not much evidence against homoscedasticity despite the substantial leptokurtosis in the data. The latter is primarily driven by the extreme observation at the end of September 1985, when the precipitous drop of the dollar was particularly large versus the yen. The evidence for a second-order VAR and for quarterly returns is qualitatively similar.

3.2 A Conditional-Volatility Model

In the framework used here, conditional volatility in the forward market depends on both the parameters of the VAR and the conditional variances of its residuals. If the residual vector of the VAR is denoted by u_{t+1}, the conditional variance is given by

$$hrp_{t,n} = e1' \sum_{i=0}^{n-1} (I - A)^{-1}(I - A^{i+1}) \times E_t[u_{t+n-i}u'_{t+n-i}](I - A^{i+1})'(I - A)^{-1'}e1. \quad (9)$$

Let $H_{t+1} = E_t[u_{t+1}u'_{t+1}]$. When the VAR is of higher order, H_{t+1} contains zeros corresponding to the lagged values in the companion form of the VAR. I model time variation in H_t with a multivariate GARCH model developed by Baba, Engle, Kraft, and Kroner (1989):

$$H_{t+1} = C'C + D'u_tu'_tD + G'H_tG, \quad (10)$$

where C, D, and G are 6×6 parameter matrices and C is symmetric. The model results in a positive-definite H_t under very weak conditions. Let vec(.) denote the VEC operator Then the model can be rewritten as

$$\text{vec}(H_{t+1}) = C^* + D^*\text{vec}(u_tu'_t) + G^*\text{vec}(H_t), \quad (11)$$

with $C^* = \text{vec}(C'C)$, $D^* = D' \otimes D'$, and $G^* = G' \otimes G'$. The conditional variance of each variable is related to past squared residuals and cross-residuals and past variances and covariances of all variables. The full model contains 93 parameters, which were estimated using maximum likelihood with a normal conditional distribution function for the innovations of the monthly second-order VAR. This can be viewed as quasi-maximum likelihood (see Bollerslev and Woolridge 1992) if "robust" standard errors are computed as we shall do in Table 11.

Table 10 reports several likelihood ratio tests that examine the importance of cross-dependencies between different currency volatilities and between exchange-rate and forward-premium volatility. I consider five restricted models. The

Table 11. Parameter Estimates From a Multivariate GARCH Model: Parameter Estimates

	$D^*_{1,1}$	$D^*_{2,8}$	$D^*_{3,15}$	$D^*_{4,22}$	$D^*_{5,29}$	$D^*_{6,36}$	$G^*_{1,1}$	$G^*_{2,8}$	$G^*_{3,15}$	$G^*_{4,22}$	$G^*_{5,29}$	$G^*_{6,36}$
$H_{i,i}$	.044	.067	.056	.034	.100	.040	.913	.824	.895	.957	.867	.963
	(.007)	(.016)	(.016)	(.017)	(.004)	(.029)	(.031)	(.056)	(.065)	(.089)	(.013)	(.141)

NOTE: Parameters of the BEKK model in Equation (10) in the text are obtained by maximum likelihood using a multivariate normal distribution for the innovations of the second-order VAR on monthly forward-market returns. Under very weak conditions including misspecification of the distribution function (see Bollerslev and Woolridge 1992; White 1982), the vector of parameters is asymptotically normally distributed with covariance matrix $\Omega = A^{-1}BA^{-1}$, where A is the Hessian form and B the outer product form of the information matrix. The reduced-form coefficients from Equation (11) are reported here. To conserve space, I only report the dynamics of the conditional variances $H_{i,i}$ ($i = 1, 2, 3, 4, 5, 6$) restricted to the coefficients on the own lagged conditional variance or squared residual.

first model is a diagonal model, restricting each conditional variance to depend only on the own past conditional variance and squared innovation. The second model in addition enters the conditional variance and squared innovation of the own forward premium (exchange-rate change) in each exchange-rate (forward-premium) volatility equation. This is a test of cross-market dependence.

The other tests examine the dependencies between exchange-rate and forward-premium volatilities. The fourth (fifth) model imposes no influences of conditional (co) variances of forward premiums (exchange rates) in exchange-rate (forward-premium) equations. The third model imposes both of these restrictions. All of these restricted models are firmly rejected. Whether this volatility predictability is due to persistent "fundamental" shocks across currencies and asset markets, policy coordination, or various types of market failures remains an open question.

Maximum likelihood estimates of the coefficients in C^*, D^*, and G^* are partially reported in Table 11. The maximum eigenvalue of $(D^* + G^*)$ is .994, whereas the smallest eigenvalue is .848. This indicates that the system is nearly integrated in variance (see Bollerslev and Engle 1993), which is symptomatic for financial series. The coefficients on own past variances, for instance, vary between .824 for the mark exchange rate and .963 for the pound forward premium. Diagnostic testing of the normalized residuals (Table 12) reveals that the BEKK model successfully eliminates most of the time variation in squared residuals of all variables, except for the pound forward premium. It turns out that it is very hard to account for all of the time variation in the cross residuals.

Using Equation (9), it is straightforward to compute hrp_t. Details of the procedure are given in an appendix available from the author. The implied volatility of monthly forward-market investments is graphed in Figure 3. Many interesting patterns emerge. The ex ante volatilities of the three returns seem to move together. For instance, sudden increases in uncertainty occurred, albeit of different magnitude, at the end of 1978 and 1981 and in 1985 across markets. The extreme peaks in 1985 coincide with the moment the dollar started its precipitous decline, which was not anticipated by the VAR

Table 12. Parameter Estimates From a Multivariate GARCH Model: Residual Diagnostics

	Q2(4)	Q2(13)	Q2(26)
$H_{1,1}$	.830	2.498	13.874
	(.934)	(.999)	(.975)
$H_{2,2}$	9.926	18.685	34.523
	(.042)	(.133)	(.122)
$H_{3,3}$	2.125	7.341	18.147
	(.713)	(.884)	(.870)
$H_{4,4}$	9.005	9.441	10.690
	(.061)	(.739)	(.996)
$H_{5,5}$	4.321	12.243	20.798
	(.364)	(.508)	(.752)
$H_{6,6}$	13.195	28.822	34.148
	(.010)	(.007)	(.131)

NOTE: Parameters of the BEKK model in Equation (10) in the text are obtained by maximum likelihood using a multivariate normal distribution for the innovations of the second-order VAR on monthly forward-market returns. Under very weak conditions including misspecification of the distribution function (see Bollerslev and Woolridge 1992; White 1982), the vector of parameters is asymptotically normally distributed with covariance matrix $\Omega = A^{-1}BA^{-1}$, where A is the Hessian form and B the outer product form of the information matrix. This table reports the Ljung–Box statistics on the squared VAR residuals normalized by the estimated conditional variances. If the normalized residuals are serially uncorrelated, the statistics are $\chi^2(n)$, where n is the number of autocorrelations included. The p values are reported in parentheses.

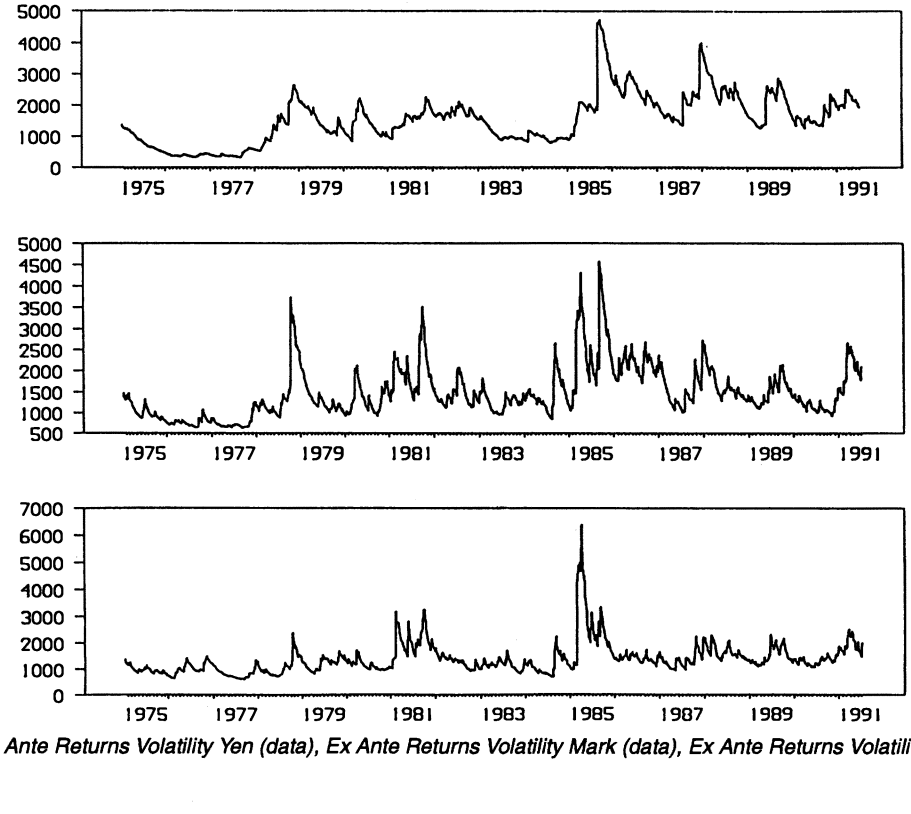

Figure 3. Ex Ante Returns Volatility Yen (data), Ex Ante Returns Volatility Mark (data), Ex Ante Returns Volatility Pound (data).

system. In general, volatility was somewhat higher during the eighties than during the seventies.

4. CONCLUSIONS

This article has studied conditional means and variances of both monthly and quarterly yen, mark, and pound forward-market returns in an integrated VAR framework using weekly data. A first set of results concerns the predictability of forward-market returns. Whereas it is well known that the forward premium predicts the corresponding future forward-market return, I find that an increase of the forward premium on the yen, mark, or pound signals lower future forward-market returns on all three currencies. Moreover, this predictability carries over to the conditional volatility of forward-market returns. Past second moments of forward premiums predict the volatility of exchange rates. Moreover, the predictability of yen/mark, yen/pound, and mark/pound returns is similar to the predictability of dollar returns. Finally, I find that the most recent weekly change in the spot exchange rate has more predictive power for future exchange-rate changes than other components of past monthly or quarterly exchange-rate changes. This result is inconsistent with the often-heard claim that exchange rates are random walks with respect to the information set generated by their past realizations.

If the predictability is due to risk, the VAR results imply that the risk premiums on forward-market investments are very variable and persistent, and they change sign quite often. The standard deviation of the risk premium on all three investments is around 10%. On average, the premiums on cross-rate investments are somewhat less variable. Using a nonlinear Wald test, the hypothesis that expected forward-market returns move perfectly together pairwise could never be rejected. The VAR's implied correlation between the yen premium and the premiums on the European currencies is substantially weaker, however, than between the European currencies themselves. Traditional latent-variable tests fail to pick up such subtle but economically significant differences in comovements of expected returns.

The implications for asset pricing are therefore both encouraging and challenging. The strong comovement between the various expected returns raises hope that a limited number of market fundamentals are the common cause of the predictability. The large variability of the risk premiums, on the other hand, poses a heavy burden on economic theory because many fundamentals—for example, aggregate consumption or asset supplies—do not show such variability. As an example, Korajczyk (1985) attempted to link the forward-market risk premium to real-interest differentials. Even if a large fraction of the variance of ex post real-interest differentials can be explained with current information variables, the results derived in this article imply that it would not suffice to account for the risk-premium variability. Similarly, models that link risk premiums to the conditional exchange-rate variance require both comovement between expected returns and conditional variances and sufficient variability in conditional variances (see Frankel 1988; Giovannini and Jorion 1989). The empirical estimates in this article yield quite variable conditional variances, but their variability is an order of magnitude less than the variability of expected returns, and they show little correlation with both expected returns and their absolute values.

Finally, the VAR framework, spelled out in this article could be fruitfully applied in other contexts. In view of the interactions between equity and foreign-exchange markets that have been recently detected (see, for instance, Bekaert and Hodrick 1992), it would be particularly interesting to combine high-frequency stock-return data with high-frequency foreign-exchange market data. There could also be interactions between conditional means and conditional variances that were ignored in the current analysis. All of this is left for future research.

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