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THE RISK PREMIUM IN FORWARD FOREIGN EXCHANGE MARKETS AND G-3 CENTRAL BANK INTERVENTION: EVIDENCE OF DAILY EFFECTS, 1985-1990

by Richard T. Baillie and William P. Osterberg

Richard T. Baillie is a professor of economics at Michigan State University, East Lansing, Michigan. William P. Osterberg is an economist at the Federal Reserve Bank of Cleveland. The authors are grateful to Ralph Day, Kyle Fleming, and Rebecca Wetmore for research assistance and to Owen Humpage for useful comments and suggestions.

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ABSTRACT

Evidence that forward rates for foreign exchange are not unbiased forecasts of future spot rates suggests a time-varying risk premium. However, there is little evidence that the forecast error is related to fundamentals, although most investigations have lacked high-frequency data. In this paper, we use daily exchange-rate and official Federal Reserve intervention data to test for an impact of intervention on the forecast error. This paper extends recent analyses of daily changes in exchange rates by Baillie and Bollersev (1989) and Hsieh (1989) to the daily forward-rate forecast errors for the dm/US$ and yen/US$ rates. We estimate an MA(21) process and utilize GARCH with a conditional student-t distribution. We find that 1) U.S. purchases of dollars on day t-1 affect the day t forecast error (f_t-$E_t[s_{t+k}]$), 2) there are day-of-the-week effects in the conditional variance, and 3) for the yen/US$ rate, there is GARCH-in-mean. These findings provide some support for considering intervention as a channel through which fundamentals influence risk premia.

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I. Introduction

The view that widely and frequently traded asset prices should reflect all available information is one of the most widely held tenets in economics and finance. Because foreign exchange markets are worldwide in scope and almost nonstop in operation, the large number of tests of exchange-market efficiency is not surprising. One of the most noteworthy findings to date has been the tendency for changes in exchange rates to be uncorrelated, but with fat-tailed distributions. There are distinct periods of high or low variance, so that volatility appears in clusters. In the case of forward markets for foreign exchange, rejection of efficiency can conceivably be explained by a risk premium, which, as we indicate below, may be related to time-varying conditional heteroscedasticity.

Interest has heightened in studying the role of central bank intervention in influencing exchange rates. During the period of ostensibly floating rates, central bank intervention policy has been designed both to influence the level of the exchange rate and to reduce its volatility. Specifically, as discussed by Funabashi (1989) and Dominguez (1990), soon after the Plaza accord in September 1985, the Group of Three (G-3) finance ministers agreed to reduce the dollar's exchange value. Then, at the Louvre meeting in 1987, they decided to shift to a regime of stabilization. Thus, there is a clear interest in analyzing the impact of intervention during this period on both the level and volatility of exchange rates, although the academic literature is undecided as to how intervention may influence either one.

This paper has two purposes: First, we seek evidence of a risk premium in the forward rate for foreign exchange. Second, we add to a growing body of literature analyzing the impact of central bank intervention during the period

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of floating rates. This paper is unique in that it analyzes the impact of intervention on the forward rate with daily data, allowing a time-varying risk premium to emerge via a GARCH (generalized autoregressive conditional heteroscedasticity) formulation in which intervention may influence the conditional variance. It is by now well known that volatility measures deteriorate with longer sample periods such as those previously employed to analyze volatility in forward markets. We avoid this outcome by using daily data. However, in the absence of observations on expected future rates, the use of daily data implies the presence of a high-order process describing the forward forecast error (see footnote 3 on page 5). This paper is organized as follows: In the next section we review the relevant literature on risk premia in forward markets and the evidence for an impact of daily intervention. In section III we present the model that is analyzed empirically. In section IV we discuss the data, and in section V we present the results of the empirical investigation. Finally, we conclude in section VI.

II. Related Literature

The conjecture that forward rates are unbiased and efficient predictors of future spot rates has been widely tested. In theory, unbiasedness holds only given rational expectations and risk-neutrality of the representative investor. Most studies have used weekly data, which imply serially correlated forecast errors because the sampling interval is then finer than the forecast interval, which is one month for a one-month-forward contract. As summarized by Baillie (1989), a consensus against unbiasedness has emerged. The possible explanations include the inappropriateness of the rational expectations assumption (see Frankel and Froot [1987]), the possibility that policy changes

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would lead to ex post biasedness even if unbiasedness held ex ante (Lewis [1988]), anticipation of real exchange-rate changes (Levine [1989]), or the existence of a time-varying risk premium. A variety of theoretical approaches, summarized by Hodrick (1987), imply a time-varying risk premium.

An early approach by Lucas (1978) relates the risk premium to the conditional covariance between a long position in the forward market and the marginal rate of substitution between future and current consumption. Hodrick (1989) shows how the risk premium in the forward market can be more directly related to the conditional variance of market fundamentals such as money supply and government spending. Domowitz and Hakkio (1985) use monthly data to test for ARCH in the forecast error and for an influence of the conditional variance on the forecast error. While they reject efficiency, they find little evidence that the forecast error is related to the conditional variance. In general, evidence in favor of the existence of a risk premium in the forward market is weak (see also Engel and Rodrigues [1989], Kaminsky and Peruga [1990], and Mark [1988]). This may partly reflect the need to use data of no higher than monthly frequency in analyzing the relationship between the forward-rate forecast error and either consumption or money.[1] Baillie and Bollersev (1989) have noted that volatility measures such as conditional variance exhibit less time variation when they are constructed from data of lower frequency.

[1]There are indirect approaches to testing for a risk premium using daily data. One approach is that taken by Levine (1989), who tests the implication of many asset pricing models that the risk premium imbedded in the forward rate is exactly equal to the risk premium in the differential in real interest rates. Giovanni and Jorion (1987) test for the influence of various proxies for a risk premium, such as lagged forward rates and squared interest rates.

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While analyses of the risk premium in the forward market have been hampered by a focus on data of relatively low frequency, recent analyses by Baillie and Bollersev (1989), Hsieh (1988), and Milhoj (1987) have supported the application of GARCH to the analysis of daily exchange-rate movements. Many studies of floating exchange-rate regimes have concluded that while both spot and forward rates appear to have unit roots in their univariate representations, their distributions are unimodal, symmetric, and fat-tailed. GARCH allows for a conditionally normal distribution that is unconditionally symmetric and leptokurtic. GARCH has been extensively utilized to model exchange-rate volatility (see Engle and Bollersev [1986], Bollersev [1987], Hsieh [1989], Diebold and Nerlove [1989], McCurdy and Morgan [1989], and Milhoj [1987]). Baillie and Bollersev (1989) modify GARCH to consider a conditionally leptokurtic distribution that is capable of accounting for severe leptokurtosis in the daily data. In a multivariate setting, Baillie and Bollersev (1990) apply GARCH to analyze the risk premium in the forward market with weekly data. They find no evidence that the forward forecast error can be explained by its conditional variance, as predicted by various theoretical approaches.

A description of the theoretical channels of influence for intervention is given by Obstfeld (1989). The portfolio balance channel is the influence of sterilized central-bank intervention on the relative magnitude of portfolios of securities denominated in different currencies. If investors are risk averse and view assets of different currency denominations as imperfect substitutes, shifts in asset supplies may induce changes in exchange rates. Most empirical investigations conclude that there is no portfolio balance effect. However, the need to calculate aggregate portfolio shares

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leads to the use of relatively low-frequency data.

A second channel of influence for intervention could be via signaling.[2] The effectiveness of intervention in this case depends on the credibility of the signal. If intervention could be signaling only future monetary policy, it is unclear why intervention would be chosen over alternative signals such as "cheap talk" (Stein [1989]). However, in this case, once the central bank has intervened, it may stand to lose money by not following through on the expected policy. Dominguez (1988) looks at weekly money supply announcements and finds evidence that the impact of intervention depends on the credibility of the implied monetary policy action.[3] In general, it is difficult to disentangle portfolio balance and signaling influences. Ghosh (1989) and Dominguez and Frankel (1989) present recent attempts to disentangle the two channels.

In this paper, we do not distinguish between the two. Our approach is closer to that of Domowitz and Hakkio (1985) and Osterberg (1989). Domowitz and Hakkio show how changes in money supplies in a two-country model of exchange-rate determination can influence the risk premium in the forward rate either by influencing the conditional mean of the forward forecast error or by influencing its conditional variance. Osterberg modifies Hodrick (1989) to

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[2]A relatively new but rapidly growing body of research views intervention as a signal that the market can use to infer target bands for exchange rates. However, the objective of this research is not to test for an impact of intervention, which is distinguished from the fundamentals that determine the equilibrium level of the exchange rate. See, among others, Froot and Obstfeld (1989) and Klein and Lewis (1991).

[3]However, if intervention is a useful signal only if the monetary authorities follow through with the expected future policy, there is an obvious difficulty in attributing any exchange-rate movement to the intervention and not to expected monetary policy. See Humpage (1991), Klein and Rosengren (1991), and Dominguez (1990) for discussions of this issue.

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show how intervention, by changing the amount of privately held currency, can influence the forward rate's risk premium through both the conditional mean of the forward-rate error and its conditional variance. We know of no studies isolating an impact of intervention on the conditional variance of the forward-rate forecast error. However, Loopesko (1984) and Dominguez (1990) find influences of intervention on the risk premium implied by the uncovered interest parity condition.

III. The Model

Equation (1) expresses the decomposition of the forward-rate error, $s_{i,t+k}-f_{i,t}$, into a risk premium $\delta_{i,t}$ and a forecast error $\mu_{i,t}$. $s_{i,t+k}$ and $f_{i,t}$ are the log of the spot rate at time t+k and the log of the forward rate at time t for a contract that settles at t+k, respectively, for currency i.

$$s_{i,t+k} - f_{i,t} = \delta_{i,t} + \mu_{i,t+k} \tag{1}$$

When the forecast horizon, k, is longer than the sample frequency, the forecast error will be autocorrelated. Specifically, the autocorrelation coefficient at lag j will equal zero only for $j \geq [k] + 1$, where [k] is the largest integer smaller than k. As we discuss below, settlement conventions in the foreign exchange markets suggest that k = 22. As discussed by Baillie (1989), the simplest model for $\mu_{i,t+k}$ is MA([k]), expressed in equation (2).

$$\mu_{i,t+k} = \sum_{j=1}^{[k]} \theta_j \epsilon_{i,t+k-j} + \epsilon_{i,t+k} \tag{2}$$

Conceivably, rather than freely estimating such a high-order MA process, we could impose coefficients suggested by theory to improve the power of our estimation of the coefficents on the variables of interest. Baillie and Bollersev (1990), in their study of weekly observations on the forward-rate

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forecast error, impose the four MA coefficients suggested by the assumption that $s_{i,t+k}$ is a martingale. Here the MA coefficients are estimated freely.[4] Equation (3) presents our model of the risk premium.

$$\delta_{i,t} = b_0 + \sum_{j=1}^{5} b_{Dj} D_{j,t} + \sum_{j=1}^{4} b_{Ij} I_{j,t-1} + b_{h,p} h^p \tag{3}$$

Equation (3) indicates that, other than a constant component, b_o, we allow $\delta_{i,t}$ to exhibit day-of-the-week effects (D) to be influenced by intervention (I) and to be related to conditional variance h (p=1,2 denote conditional variance and standard deviation). We cannot hope to distinguish between inefficiency and risk because the significance of the coefficients on intervention could simply reflect the failure of the market to take account of available information. Details about these variables are given below. Equations (2) and (3) are combined to yield equation (4).

$$s_{i,t+k} - f_{i,t} = b_0 + \sum_{j=1}^{21} \theta_j \epsilon_{i,t+k-j} + \sum_{j=1}^{5} b_{Dj} D_{j,t} + \sum_{j=1}^{4} b_{Ij} I_{j,t-1} + \epsilon_{i,t+k} + b_{h,p} h^p,\ p=1,2 \tag{4}$$

Conditional normality of the errors implies an unconditional, symmetric, but fat-tailed distribution. However, we allow in equation (5) for a conditional student-t distribution that may be more successful in explaining leptokurtosis (see Baillie and Bollersev [1989] and Hsieh [1989]).

$$\epsilon_t | \psi_{t-1} \sim t(0, h_t^2, \upsilon) \tag{5}$$

As the distributional parameter, υ, approaches 30, this distribution is

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[4] In fact, in our model, if we were to impose the martingale assumption, we would imbed strict noninvertibility (Harvey [1981]) into the system. The reason is not that we are analyzing daily data, but rather that the forecast interval, k, is an exact integer multiple of the sampling frequency. In studies of the forecast error utilizing weekly data, k = 4-2/5. See Baillie and Bollersev (1990).

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close to normal. Equation (6) indicates that the conditional variance is modeled as a GARCH(1,1) with intercept and the possibility of impacts for daily dummies and central bank intervention.

$$h_t^2 = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1}^2 + \sum_{j=1}^{5}\gamma_{D,j}\,D_j + \sum_{j=1}^{4}\gamma_{I,j}\,I_{j,t-1} \tag{6}$$

Equations (4) and (6) will be estimated simultaneously.

IV. Data

The exchange rate data were provided by the Federal Reserve Bank of New York. At 10:00 a.m. of each day on which the New York market is open, the Bank obtains both bid and ask quotes for the spot rates and forward premium $(s_t - f_t)$. The forward rate is thus calculated as simply the spot rate plus the premium. Although some authors have averaged bid and ask quotes, we use only the bids. Bossaert and Hillion (1991) show why averaging is inappropriate, presenting evidence that previous conclusions on the efficiency of the forward market may be reversed when only bids or asks are analyzed. They also claim that intervention may impact the bid-ask spread.

We match the spot and forward rates so that s_{t+k} and f_t are quotes on contracts that settle on the same day. Riehl and Rodriguez (1977) describe the mechanics of contract settlement in the foreign exchange market, which are essentially as follows: Find the day on which the contract corresponding to f_t would settle, go forward two business days, then go forward to the same day in the next month. If that day is not a business day, go forward until one is found, unless this implies a day in the third month, in which case the last business day in the month is chosen. The future spot rate, s_{t+k}, that settles on the same day is the one quoted two business days prior to the day on which

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the forward contract settles. This allows for settlement of the spot and forward contract on the same day. Levine (1989) provides evidence that failure to match spot and forward rates correctly may have influenced previous findings on the extent to which the forward forecast error is influenced by the risk premium presumably imbedded in real interest-rate differentials.

The intervention data were provided to us by the Board of Governors of the Federal Reserve System. For each G-3 country, we utilize the actual amount of net daily dollar purchases. This enables us to avoid the pitfall of introducing simultaneity through the conversion from a raw foreign currency magnitude to dollars through application of the exchange rate. The data are close-of-business (COB) amounts. Thus, we align the 10:00 a.m. quotes on day t with intervention dated t-1, which is the net intervention from COB t-2 to COB t-1. Since intervention may occur on holidays, we add such intervention to the previous day's amount. In other words, if Monday is a holiday on which there is intervention, the Tuesday 10:00 a.m. rate quote is aligned with the total of net intervention from COB Thursday to COB Friday (the original Friday number) plus that occurring on Monday.[5]

We also transform the intervention data so that the coefficent on intervention can be interpreted as the elasticity of the premium with respect to intervention. Since negative intervention observations represent sales of dollars (purchases of yen or Deustche marks), for each bilateral relationship we have four intervention measures (I_js): both purchases of dollars and dollar

[5]The dummies are constructed so as to be orthogonal. In other words, if Friday is a holiday, the holiday dummy for the Monday observation has a value of one, but the Monday dummy does not. In order to avoid the dummy variable trap with the presence of a constant term, we omit one of the daily dummies. Most holidays fall on Monday, so it is natural to omit Tuesday. All of the dummies are aligned with the day of the forward quote, rather than with the day on which the future spot quote is taken or with the day on which both settle.

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sales for each country. The intervention variable is $\max(0,\ln[I_j])$. Interventions are measured in hundred-million-dollar units and never lie in the interval (0,1].

We do not control for the influence of intervention on expected monetary policy by including data on expectations. Assuming that all interventions are sterilized, our intervention measures correspond to the dollar value of the changes in private bond portfolios resulting from the sterilizations. A truer measure of shifts in portfolio balance would be obtained if only the net change in relative portfolios were measured and if we did not distinguish between sales and purchases or identify the central bank's country.

V. Results

We utilize the Berndt et al. (1974) algorithm to obtain maximum likelihood estimates of the basic models for both the DM/US$ and yen/US$ models given by equations (4), (5), and (6). These results are presented in table 1. Though the sample periods are the same for both currencies, the sample sizes differ due to a dissimilar number of market holidays for Japan and West Germany. Columns (a) and (d) are for simple models with constants in both mean and variance, the MA(21) error structure for the mean, and conditional normality for the variance. ν is fixed at a high enough value that the t-distribution specified in equation (5) is approximately normal. All 21 MA coefficients are significant at the 5 percent level. We also calculate the roots of the lag operator polynomials implied by the estimated MA coefficients and find that they are consistent with invertibility, lying outside the unit

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circle.[6] We report the Ljung-Box (1978) statistics for kth order serial correlation in the squared residuals, $Q^2(k)$, which, under conditional homoscedasticity, are distributed as chi-squared with k degrees of freedom. For both simple models, these statistics are significant at the 5 percent level. m_3 and m_4 are distributed as N(0,6/NOBS) and N(0,24/NOBS), respectively, under normality. For both models there is significant kurtosis, suggesting that we try an assumption other than conditional normality.

Columns (b) and (e) then retain the assumption of conditional normality but model the conditional variances as GARCH(1,1) processes. The additional parameters are significant both individually and in terms of the reduction in log of the likelihood (Log-L). The Q^2 statistics are reduced, though still significant at 5 percent. Columns (c) and (f) relax the assumption of conditional normality. The values of $1/\gamma$ are obtained from iterating on γ until the approximation $m_4 = 3(\gamma-2)/(\gamma-4)$ holds. Significant reductions in Log-L are obtained with these distributions. Although examination of the Q^2 and m_4 statistics does not confirm a reduction in kurtosis, the parameterizations of columns (c) and (f) are maintained for subsequent estimations in which the values of γ are held constant.

Table 2 indicates the results of likelihood ratio tests for the inclusion of daily dummies (including a holiday dummy) or intervention in either the mean or variance equations. Given previous research utilizing GARCH to study daily exchange rates, we may expect day-of-the-week effects to be present in the mean. However, the first line indicates that there are no

[6]These calculations were performed using the GAUSS routine "polyroot," which yields a vector of 22 values in the form a +/- b*i*. We then calculate and examine the elements in the vector of $(a^2 + b^2)^{1/2}$. These results are available from the authors.

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such effects for either currency. Assuming efficiency, a significant influence of intervention in the mean could be interpreted as support for the influence of intervention via a risk premium. The second line shows that the intervention variables are not jointly significant for either currency. Since intervention is sometimes coordinated, however, collinearity may exist among the individual intervention variables, and we thus test for their significance. Only U.S. purchases of U.S. dollars (b_{I3}) have significant (positive) influences on the forward forecast error for either the DM/US$ or yen/US$ model. We retain these variables in subsequent specifications of the variance equation.

Adding all dummies to the variance equation (γ_{Dj}s) contributes significantly to both models. However, any possibility that intervention influences a risk premium via GARCH-in-mean is ruled out by the insignificance of the impact of intervention on the conditional variance (γ_{Ij}s). Last, we test for the presence of GARCH-in-mean where either the conditional variance (h=1) or the conditional standard deviation (h=2) enters the mean equation. For both specifications, we find significant effects for the yen/US$ model but not for the DM/US$ model.

VI. Summary

This paper has two somewhat disparate purposes. Using forward and spot exchange-rate data correctly matched for both the DM/US$ and yen/US$ models, we have 1) extended the GARCH with student-t parameterization to the daily forward-rate forecast error in an attempt to find evidence of a time-varying risk premium and 2) looked for an effect of intervention in the daily forward market. Previous investigations of the forward-rate error have used lower-

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frequency data, which reduce measured volatility and thus affect the chances of finding GARCH-in-mean, one of the channels through which fundamentals may influence risk premia. Intervention is one of the few variables measured at a daily frequency that could be considered fundamental.

We model the forecast error as an MA(21) process and find that the student-t parameterization is a significant improvement over conditional normality. This is similar to the findings of previous research on the daily exchange-rate process. Our evidence of a time-varying risk premium is mixed. For the yen/US$ rate, we find GARCH-in-mean, but no influence of intervention on the conditional variance. For both currencies, we find an influence of U.S. purchases of dollars on the conditional mean of the forecast error.

We cannot claim to have distinguished between the signaling and portfolio balance channels. However, if the portfolio balance channel were operative, we would expect that purchases and sales would have equal influence and that the nationality of the central bank would make no difference. We have not tested these hypotheses at this point.

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TABLE 1: Parameter Estimates for the Basic Model

$$(4A)\ s_{i,t+k} - f_{i,t} = b_0 + \sum_{j=1}^{21} \theta_j \epsilon_{i,t+k-j} + \epsilon_{i,t+k} \quad (6A)\ h_t^2 = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1}^2$$

	DM/US$			Yen/US$		
	(a)	(b)	(c)	(d)	(e)	(f)
b_0	-0.6252	-0.6346	-0.4385	-0.3705	-0.2405	0.0839
	0.4384**	0.4173	0.4103	0.4200	0.3890	0.3234
θ_1	0.9324	0.9339	0.9331	0.9669	0.9767	0.9558
	0.0303**	0.0316**	0.0309**	0.0294**	0.0337**	0.0294**
θ_2	0.8933	0.8758	0.8787	0.9773	0.9800	0.9419
	0.0378**	0.0435**	0.0419**	0.0409**	0.0487**	0.0419**
θ_3	0.8833	0.8634	0.8690	0.9541	0.9417	0.9338
	0.0451**	0.0497**	0.0478**	0.0510**	0.0607**	0.0503**
θ_4	0.8538	0.8373	0.8449	0.9149	0.8958	0.8868
	0.0485**	0.0523**	0.0520**	0.0559**	0.0644**	0.0553**
θ_5	0.8553	0.8333	0.8296	0.8916	0.8607	0.8461
	0.0535**	0.0566**	0.0552**	0.0577**	0.0641**	0.0570**
θ_6	0.8406	0.8318	0.8256	0.8505	0.8214	0.8042
	0.0535**	0.0574**	0.0560**	0.0585**	0.0659**	0.0586**
θ_7	0.8322	0.8127	0.8113	0.8396	0.8185	0.8003
	0.0555**	0.0605**	0.0578**	0.0609**	0.0658**	0.0582**
θ_8	0.8209	0.7989	0.7945	0.8491	0.8187	0.8097
	0.0575**	0.0609**	0.0588**	0.0621**	0.0658**	0.0597**
θ_9	0.8413	0.8228	0.8136	0.8227	0.8072	0.8126
	0.0587**	0.0605**	0.0587**	0.0647**	0.0679**	0.0611**
θ_{10}	0.8428	0.8309	0.8332	0.8311	0.8133	0.7959
	0.0576**	0.0598**	0.0586**	0.0639**	0.0662**	0.0620**
θ_{11}	0.7529	0.7558	0.7670	0.8133	0.7813	0.7736
	0.0582**	0.0602**	0.0585**	0.0632**	0.0681**	0.0633**
θ_{12}	0.7883	0.7757	0.7851	0.8202	0.7999	0.7703
	0.0562**	0.0593**	0.0576**	0.0606**	0.0639**	0.0615**
θ_{13}	0.7702	0.7562	0.7553	0.7414	0.7292	0.6708
	0.0577**	0.0598**	0.0582**	0.0634**	0.0668**	0.0607**
θ_{14}	0.7577	0.7489	0.7481	0.7649	0.7568	0.7133
	0.0589**	0.0612**	0.0588**	0.0647**	0.0682**	0.0601**

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Table 1 (continued)						
θ_{15}	0.7538	0.7374	0.7462	0.7733	0.7660	0.7336
	0.0578**	0.0596**	0.0565**	0.0645**	0.0678**	0.0582**
θ_{16}	0.6856	0.6754	0.6876	0.7419	0.7146	0.6690
	0.0529**	0.0535**	0.0528**	0.0634**	0.0687**	0.0581**
θ_{17}	0.6284	0.6178	0.6182	0.6707	0.6729	0.6440
	0.0512**	0.0523**	0.0518**	0.0605**	0.0673**	0.0555**
θ_{18}	0.6417	0.6277	0.6139	0.5419	0.5557	0.5134
	0.0509**	0.0515**	0.0498**	0.0581**	0.0626**	0.0520**
θ_{19}	0.5101	0.5045	0.4902	0.4822	0.4715	0.4078
	0.0486**	0.0495**	0.0465**	0.0490**	0.0529**	0.0454**
θ_{20}	0.3856	0.3760	0.3587	0.2797	0.2735	0.2218
	0.0422**	0.0436**	0.0407**	0.0398**	0.0425**	0.0377**
θ_{21}	0.1652	0.1587	0.1517	0.1119	0.1029	0.0825
	0.0313**	0.0319**	0.0301**	0.0304**	0.0327**	0.0284**
ω	0.9227	0.0317	0.0308	0.8876	0.0205	0.0204
	0.0324**	0.0138**	0.0166*	0.0266**	0.0068**	0.0095**
α		0.0515	0.0497		0.0386	0.0522
		0.0149**	0.0169**		0.0064**	0.0135**
β		0.9133	0.9159		0.9375	0.9198
		0.0263**	0.0303**		0.0120**	0.0213**
$1/\gamma$	0.01-fixed	0.01-fixed	0.0995-fixed		0.01-fixed	0.1710-fixed
Log-L	-1535.4580	-1513.9236	-1502.4570	-1489.4215	-1462.0421	-1402.7537
Q(20)	1.4187	3.5811	4.1893	4.3272	9.7477	16.0533
$Q^2(20)$	105.3799	46.2830	46.7521	72.7154	33.0963	29.6398
m_3(skewness)	-0.1636	-0.0801	-0.0772	-0.3538	-0.3072	-0.3398
m_4(kurtosis)	4.3781	3.9066	3.9919	5.6407	5.7382	6.2488
$3(\gamma-2)/(\gamma-4)$	NA	NA	3.9917	NA	NA	6.2486
NOBS	1113	1113	1113	1095	1095	1095

* Significant at the 5 percent level.
** Significant at the 10 percent level.
NOTE: Standard errors are beneath coefficient estimates.
SOURCE: Authors' calculations.

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TABLE 2: Likelihood Ratio Tests for Model Specification		
	DM/US$	Yen/US$
$b_{Dj} = 0$, j=1,2,3,4,5.	1.552	3.281
$b_{Ij} = 0$, j=1,2,3,4.	3.262	5.410
$\gamma_{Dj} = 0$, j=1,2,3,4,5.	27.025**	14.604**
$\gamma_{Ij} = 0$, j=1,2,3,4.	0.390	4.825
$b_{I3} = 0$, j=3 denotes U.S. purchases	2.783*	3.521*
$b_{I3} \neq 0$, $\gamma_{Dj} = 0$, j=1,2,3,4,5.	27.265**	15.852**
$b_{I3} \neq 0$, $\gamma_{Ij} = 0$, j=1,2,3,4.	0.310	3.108
$b_{i3} \neq 0$, $\gamma_{Dj} \neq 0$, j=1,2,3,4,5, $b_{h,1}$=0.	0.017	9.272**
$b_{I3} \neq 0$, $\gamma_{Dj} \neq 0$, j=1,2,3,4,5 $b_{h,2}$=0.	0.029	9.299**

* Significant at the 5 percent level.
** Significant at the 10 percent level.
SOURCE: Authors' calculations.

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References

Baillie, Richard T., 1989, "Econometric Tests of Rationality and Market Efficiency," Econometric Reviews, 8, 151-186.

________, and Tim Bollersev, 1989, "The Message in Daily Exchange Rates: A Conditional Variance Tale," Journal of Business and Economic Statistics, 7, 297-305.

________, and ________, 1990, "A Multivariate Generalized ARCH Approach to Modeling Risk Premia in Forward Foreign Exchange Rate Markets," Journal of International Money and Finance, 9, 309-324.

Berndt, E.K., B.H. Hall, R.E. Hall, and J.A. Hausman, 1974, "Estimation and Inference in Nonlinear Structural Models," Annals of Economic and Social Measurement, 4, 653-665.

Bollersev, Tim, 1987, "A Conditional Heteroscedastic Time Series Model for Speculative Prices and Rates of Return," Review of Economics and Statistics, 69, 524-547.

Bossaert, P., and P. Hillion, 1991, "Market Microstructure Effects of Government Intervention in the Foreign Exchange Market," Review of Financial Studies, forthcoming.

Diebold, Francis X., and Marc Nerlove, 1989, "The Dynamics of Exchange Rate Volatility: A Multivariate Latent Factor ARCH Model," Journal of Applied Econometrics, 4, 1-21.

Dominguez, K.,1988, "The Informational Role of Official Foreign Exchange Intervention: An Empirical Investigation," Working Paper, Harvard University, November.

________, 1990, "Market Responses to Coordinated Central Bank Intervention," Carnegie-Rochester Conference Series on Public Policy, 32, 121-164.

________, and Jeffrey Frankel, 1989, "Have Recent Central Bank Foreign Exchange Intervention Operations Influenced the Yen?" unpublished manuscript, Harvard University and the National Bureau of Economic Research.

Domowitz, I., and C.S. Hakkio, 1985, "Conditional Variance and the Risk Premium in the Foreign Exchange Market," Journal of International Economics, 19, 47-66.

Engel, C., and A.P. Rodrigues, 1989, "Tests of International CAPM with Time Varying Covariances," Journal of Applied Econometrics, 4, 119-138.

Engle, Robert F., and Tim Bollersev, 1986, "Modelling the Persistence of Conditional Variance," Econometric Reviews, 5, 1-50.

www.clevelandfed.org/research/workpaper/index.cfm

Frankel, Jeffrey A., and Kenneth Froot, 1987, "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations," American Economic Review, 77, 133-153.

Froot, Kenneth A., and Maurice Obstfeld, 1989, "Exchange Rate Dynamics under Stochastic Regime Shifts: A Unified Approach," Harvard Institute of Economic Research Discussion Paper No. 1451.

Funabashi, Yoichi, 1989, Managing the Dollar: From the Plaza to the Louvre, 2nd Edition, Washington, D.C.: Institute for International Economics.

Ghosh, A.R., 1989, "Is It Signalling? Exchange Intervention and the Dollar-Deutschemark Rate," working paper, Princeton University, Princeton, N.J.

Giovanni, A., and P. Jorion, 1987, "Interest Rates and Risk Premia in the Stock Market and in the Foreign Exchange Market," Journal of International Money and Finance, 6, 107-124.

Harvey, A.C., 1981, Time Series Models, Oxford: Philip Allan Publishers Limited.

Hodrick, R.J., 1987, Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets, Chur, Switzerland: Harwood Academic Publishers.

________, 1989, "Risk, Uncertainty, and Exchange Rates," Journal of Monetary Economics, 23, 433-459.

Hsieh, D.A., 1988, "The Statistical Properties of Daily Foreign Exchange Rates," Journal of International Economics, 24, 129-145.

________, 1989, "Modeling Heteroskedasticity in Daily Foreign-Exchange Rates: 1974-1983," Journal of Business and Economic Statistics, 7, 307-317.

Humpage, O.F., 1991, "Central-Bank Intervention: Recent Literature, Continuing Controversy," Federal Reserve Bank of Cleveland, Economic Review, Quarter 2, 12-26.

Kaminsky, G., and R. Peruga, 1990, "Can a Time Varying Risk Premium Explain Excess Returns in the Forward Market for Foreign Exchange?" Journal of International Economics, 28, 47-70.

Klein, Michael W., and Karen Lewis, 1991, "Learning About Intervention Target Zones," manuscript.

Klein, Michael W., and Eric Rosengren, 1991, "Foreign Exchange Intervention as a Signal of Monetary Policy," New England Economic Review, May/June, 39-50.

Levine, R., 1989, "The Pricing of Forward Contracts," Journal of International Money and Finance, 8, 163-179.

www.clevelandfed.org/research/workpaper/index.cfm

Lewis, K.K., 1988, "The Persistence of the 'Peso Problem' When Policy Is Noisy," Journal of International Money and Finance, 7, 5-21.

Ljung, G.M., and Box, G.E.P., 1978, "On a Measure of Lack of Fit in Time Series Models, Biometrika, 65, 297-303.

Loopesko, B.E., 1984, "Relationships among Exchange Rates, Intervention, and Interest Rates: An Empirical Investigation," Journal of International Money and Finance, 3, 257-277.

Lucas, R.E., 1978, "Asset Prices in an Exchange Economy," Econometrica, 46, 1429-1445.

Mark, N.C., 1988, "Time Varying Betas and Risk Premia in the Pricing of Forward Foreign Exchange Contracts," Journal of Financial Economics, 22, 335-354.

McCurdy, T., and G.I. Morgan, 1989, "Evidence of Risk Premia in Foreign Currency Futures Markets, " Queens University, Department of Economics, manuscript.

Milhoj, A., 1987, "A Conditional Variance Model for Daily Deviations of an Exchange Rate," Journal of Business and Economic Statistics, 5, 99-103.

Obstfeld, Maurice, 1989, "The Effectiveness of Foreign-Exchange Intervention: Recent Experience," Working Paper No. 2796, National Bureau of Economic Research.

Osterberg, W.P., 1989, "Intervention and the Risk Premium in Foreign Exchange Rates," Working Paper 8908, Federal Reserve Bank of Cleveland.

Riehl, H., and R. Rodriguez, 1977, Foreign Exchange Markets, New York: McGraw-Hill.

Stein, J.C., 1989, "Cheap Talk and the Fed: A Theory of Imprecise Policy Announcements," American Economic Review, 79, 32-42.