Working Paper 9207

NEW RESULTS ON THE IMPACT OF CENTRAL-BANK INTERVENTION ON DEVIATIONS FROM UNCOVERED INTEREST PARITY

by Owen F. Humpage and William P. Osterberg

Owen F. Humpage is an economic advisor and William P. Osterberg is an economist at the Federal Reserve Bank of Cleveland. The authors are grateful to Richard Baillie for helpful comments and suggestions. They also thank Ralph Day, Diana Dumitru, Kyle Fleming, and Rebecca Wetmore for valuable research assistance.

April 1992

ABSTRACT

Germany, Japan, and the United States continue to view foreign exchange intervention as an effective instrument, although the mechanism through which it operates is unclear. In this paper, we use official data on daily dollar intervention to examine its impact on exchange-rate risk premia through both the portfolio-balance and expectations channels. We define the risk premium in terms of deviation from uncovered interest parity and model its behavior using generalized autoregressive conditional heteroscedasticity. Our evidence of portfolio-balance and expectations effects is inconsistent across subperiods of different exchange-rate-policy regimes. Also, unlike Dominguez (1990) and Loopesko (1984), we find no evidence that coordination of intervention improves its efficacy.

Introduction

Most large industrial countries continue to regard exchange-market intervention as a viable policy for influencing exchange rates independently of their monetary policies. In theory, sterilized intervention can affect exchange rates either through its impact on relative asset supplies or by altering expectations. Because these channels differ in their policy implications, recent studies have attempted to distinguish between them. If intervention operates by altering expectations, it is probably because the central bank has inside information about monetary policy that it is able to communicate to the market. The credibility of the implied change in policy may be crucial. If the portfolio-balance channel is operative, the shift in relative supplies of assets denominated in different currencies is all that matters.

In this paper, we use official daily intervention data to investigate the effects of German, Japanese, and U.S. (G3) intervention on exchange-rate risk premia, defined as deviations from uncovered interest parity (UIP). Loopesko (1984) and Dominguez (1990) also look at deviations from UIP, but they use different measures of intervention. Loopesko calculates cumulative intervention from the beginning of her sample, whereas Dominguez utilizes day-to-day intervention flows. We adopt both approaches here to facilitate comparisons of our results, and we also extend the analysis in a number of important ways. First, we estimate the model over a longer time frame, dividing the sample into subperiods that correspond to different U.S. intervention policies. Second, we examine the impact of dollar intervention by all other Group-of-Ten (G10) countries. Third, we control for day-of-the-week and holiday effects, which have been shown to be important in daily exchange-rate studies (see Hsieh [1988] and Baillie and Bollerslev [1989]). Finally, because of recent advances in modeling asset returns, we utilize the generalized autoregressive conditional heteroscedasticity (GARCH) framework, which allows movements in conditional variance to be analyzed directly. In addition to accommodating heteroscedastic errors, our GARCH model accounts for leptokurtosis by assuming a fat-tailed, student-t distribution.

Ideally, we would test for portfolio-balance effects by including measures of daily relative asset supplies. However, since high-quality daily data are unavailable, we use cumulative dollar intervention, which corresponds to the net dollar change in portfolios. To test for the presence of an expectations channel, we include daily intervention flows, which correspond more closely to the information provided by the most recent interventions. We also distinguish between coordinated and unilateral intervention. If just the portfolio-balance channel were operative, neither coordination nor the identities of the intervening countries would matter.

For the full sample, we find that 1) cumulative bilateral intervention influences the conditional mean of the deviation from UIP (our risk premium) for the mark-dollar case, but not for the yen-dollar case, and 2) among the categories of coordinated and unilateral interventions, only unilateral Japanese and German interventions are significant. We also find evidence of day-of-the-week effects in both the conditional mean and conditional variance. Surprisingly, when we attempt to control for regime shifts by examining subperiods, the links between intervention and risk premia are not strengthened.

The remainder of this paper is organized as follows: Section I presents a review of the literature on risk premia and intervention as background for the empirical work that follows. In section II, we present our measure of the risk premium, introduce the GARCH framework, and define alternative measures of intervention. In section III, we present our results, and in section IV, we interpret our findings.

I. Related Literature

Risk Premia in Foreign Exchange Rates

There is no consensus regarding the appropriate theoretical framework for analyzing exchange-rate risk premia. Lucas' (1982) intertemporal, dynamic, two-country model implies that risk premia should be related both to preferences and to the stochastic behavior of the driving processes, such as monetary policy. Building on this model, Hodrick (1989) relates the forward premium to conditional means and variances of market fundamentals. Osterberg (1989) extends this model further to include intervention. The intertemporal capital asset pricing model (Engel and Rodrigues [1987], Giovannini and Jorion [1989], and Mark [1988]) suggests that risk premia should be related to covariances among asset returns. The consumption-based capital asset pricing model (Hodrick [1989], Cumby [1988]) has specific implications for covariation between asset returns and intertemporal marginal rates of substitution in utility. Option pricing theory implies that risk premia are embedded in foreign-currency options prices (Lyons [1988], McCurdy and Morgan [1988]).

Tests of all of these approaches have had mixed results.[1]

Empirical evidence favoring the existence of a risk premium in foreign exchange rates is indirect. Violation of UIP and rejection of unbiasedness in the forward market both suggest that a risk premium may exist. Unfortunately, tests of UIP or of the relationship between forward and future spot rates are joint examinations of market efficiency, perfect substitutability, and capital mobility.

The poor out-of-sample forecasting performance of exchange-rate models also may reflect the existence of risk premia. The variance of exchange rates seems to show persistence, with distinct periods of low and high volatility. Such evidence has led to attempts to explain time variation in the conditional variance of exchange rates in terms of exogenous processes such as money and output. Moreover, various researchers have implied that policy shifts may be related to volatility in asset prices. Such questions about exogenous processes and policy shifts provide a motive for examining the impact of intervention on the risk premium.

Many efforts to model the conditional variances of exchange rates utilize autoregressive conditional heteroscedasticity (ARCH) or one of its variants, GARCH or and GARCH-M (see Baillie and Bollerslev [1989] and Hsieh [1989]).[2] GARCH allows for conditional normality combined with a leptokurtotic, symmetric, unconditional distribution, which is consistent with the fat tails typically found in asset-return data. Baillie and Bollerslev

---

1 Hodrick (1987) and Baillie and McMahon (1989) provide excellent overviews of this literature.

2 Domowitz and Hakkio (1985) and Baillie and Osterberg (1991) model the risk premium, defined in terms of forward forecast errors, using variants of GARCH.

(1989) successfully model the heteroscedasticity in the log first-differences of exchange rates using a version of GARCH in which the conditional distribution is a student-t. Hsieh (1989) examines distributional assumptions other than the student-t to show how they can be combined with GARCH to remove heteroscedasticity from similar data. Lastrapes (1989) uses GARCH to examine how policy regime shifts may help to explain exchange-rate volatility.

Pagan and Hong (1988) and Nelson (1987) point out the limitations of ARCH as a vehicle for explaining conditional variance, while Hodrick (1987, p. 110) argues that ARCH may be inappropriate for analyzing volatility in exchange rates. If high-risk premia are rooted in policy uncertainty, then clarification by policymakers should reduce such premia. However, because ARCH implies persistence in conditional variance, the implied risk premia would be reduced only after a period of lower ex-post volatility.

Channels of Influence for Central-Bank Intervention

Official exchange-market intervention creates an incipient change in a nation's money stock, which most large countries claim to sterilize through open-market transactions. Although sterilized intervention does not affect the monetary base, it does alter the relative stocks of domestic and foreign-currency-denominated government debt. According to the portfolio-balance theory of exchange-rate determination, if Ricardian equivalence does not hold, if bonds are imperfect substitutes in investors' portfolios, and if capital is perfectly mobile, intervention can alter exchange rates by changing the relative risk premia on governments' debt (see Weber [1986] and Backus and Kehoe [1988]). Surveying the related empirical literature, Edison (1990) finds either that there is no statistically significant relationship between risk premia and asset stocks or intervention, or that the elasticity is too small to be of any practical importance to policymakers.

Intervention may also alter exchange rates by influencing market expectations. In one version of the expectations channel, markets are viewed as weak-form efficient, and central banks can credibly signal inside information about future monetary policy via sterilized intervention (see Obstfeld [1989] and Dominguez [1988]). Klein and Rosengren (1991) find that intervention from 1985 to 1987 was instead used to clarify imprecise policy announcements. In another version of the expectations channel, exchange markets are viewed as subject to frequent, but temporary, periods of inefficiency. Investigations of survey data by Frankel and Froot (1987) and Ito (1990) tend to support this finding by casting doubt on the wisdom of applying the rational expectations assumption to exchange rates. Under such circumstances, a well-informed central bank might influence the exchange rate through intervention (see Hung [1991a]).

Some empirical studies have suggested that intervention could affect market expectations. Hung (1991b), for example, finds evidence that intervention influences unanticipated exchange-rate changes (calculated from survey data) and unanticipated volatility (calculated from option prices). Humpage (1988) shows that delineating between the first and subsequent interventions is often important, and Loopesko (1984) and Dominguez (1990) find that distinguishing between cooperative and unilateral intervention is also worthwhile.

Because there is no widely accepted theoretical model of the intervention/risk-premium mechanism, most empirical studies of intervention do not claim to have isolated the specific channels of influence. Dominguez and Frankel (1990), however, estimate both effects simultaneously and find a statistically significant impact through the portfolio-balance channel, albeit one too small to be of practical significance unless augmented by an expectations effect. Ghosh (1989) first controls for any signaling effect and then finds a significant role for portfolio balance, although one that implies a high cost for intervention.

This paper follows the approaches found in Loopesko (1984) and Dominguez (1990). Both authors examine the impact of intervention on the exchange-rate risk premia implied by deviation from UIP. Loopesko uses a cumulative measure of intervention to test for its influence through the portfolio-balance channel. Using data from 1978 through 1980, she finds mixed support for such an effect, but also shows that the coefficients on lagged exchange-rate changes and arbitrage profits are often significant, suggesting imperfectly efficient markets. Dominguez (1990) uses daily intervention flows to study the 1985-87 period and finds that the influence of intervention varies greatly over time. Both studies, however, show that coordinated intervention is more consistently significant than unilateral intervention.

II. The Empirical Model

UIP and the Risk Premium

Other studies have defined the exchange-rate risk premium in terms of forward forecast errors (see Hodrick [1987]). However, with daily data, overlapping forecast intervals resulting from this approach induce high-order serial correlation.[3] To avoid this problem, we use UIP to generate daily observations on an exchange-rate risk premium:

$$(1) \qquad RET_t = (1+R_t) - (1+R_t^*)(S_t/E_t[S_{t+1}]),$$

where

R_t = the domestic interest rate,

R_t^* = the foreign interest rate,

S_t = the exchange rate in foreign currency per U.S. dollar,

$E_t[S_{t+1}]$ = the expected one-period-ahead spot rate, and

RET_t = the excess return.

If RET_t equals zero, UIP holds. To generate an empirical measure of RET_t, we need to account for timing conventions in the foreign exchange markets. In addition, we note that in uncovered arbitrage, investors do not cover the transactions via the forward market.

Timing conventions in the spot foreign exchange markets allow for two business days between the contract date and the deliver date (see Riehl and Rodriguez [1977]). Consequently, foreign currency transactions must be completed prior to the investment. Consider an investor who expects to invest dollars overnight on day t. This investor could buy foreign currency on day t-2 for delivery on day t, invest the foreign currency overnight at date t, and receive dollars on date t+1, having sold the foreign currency proceeds on day t-1 for dollars delivered on day t+1. As part of this process, the

---

3 Baillie and Osterberg (1991) use daily forward-rate data in a GARCH framework, directly estimating the MA(21) process for the forecast error.

investor forms expectations of the uncertain future spot rate ($E_t[S_{t+1}]$) for a one-day-ahead investment. These timing conventions, together with the assumption that

(2) $E_t[S_{t+1}] = S_{t+1}$,

imply

(3) $RET_t = (1+R_t) - (1+R_t^*)(S_{t-2})/S_{t-1}) = RP_t + FE_t$.

In equation (3), we decompose the excess return into a risk premium (RP) and a forecast error (FE). By utilizing S_{t+1} instead of its expectation, we introduce an MA(1) term into FE_t. A regression of RET_t on variables in the investor's information set at the transaction time (t-2) provides a joint test of informational efficiency and of the existence of a risk premium. Hence, assuming informational efficiency, if our measure of intervention at t-3 explains RET_t, we have evidence that intervention influences the risk premium.

The Statistical Model

A substantial body of literature suggests that exchange rates are well described as martingales and that their first differences are heteroscedastic. The GARCH framework has been utilized to analyze the conditional means and variances of exchange rates with some success. In particular, the usefulness of the conditional student-t distribution in examining daily exchange-rate data has been demonstrated by Hsieh (1989) and Baillie and Bollerslev (1989). Here, we apply this framework to analyze the conditional mean and variance of the deviation from UIP. Equations (3) through (6) present our GARCH model:

(3) $$Y_t = X_t b + u_t,$$

(4) $$u_t = e_t - \theta e_{t-1},$$

(5) $$e_t | I_{t-1} \sim t(0, h_t, v),$$

(6) $$h_t = \omega + \alpha e_{t-1} + \beta h_{t-1}.$$

In equation (3), Y_t is the measured excess return (RET_t) and X_t is the vector of explanatory variables, which includes intervention, an intercept, a holiday dummy, and day-of-the-week dummies. In equation (4), the forecast error u_t follows an MA(1) process. Although ARMA analysis of RET_t does not help us to distinguish between AR(1) and MA(1) representations for equation (4), the theory implies that overlapping forecast intervals result in an MA(1) form, with which we proceed. Augmented Dickey-Fuller tests reject the hypothesis of a unit root in Y_t. Equation (5) indicates that the distribution of e_t, conditional on the information set I_{t-1}, is assumed to be student-t with a mean of zero, a variance of h_t, and a distributional parameter, v. As v approaches 30, this distribution approaches normality. As equation (6) shows, we utilize a GARCH(1,1) parameterization with an intercept.[4]

We introduce intervention (dated t-3) into the model in two forms. Following Loopesko (1984), we cumulate intervention from the beginning of our series in 1977.[5] Intervention calculated in this manner seems more

---

4 In two earlier drafts of this paper, we also explored a model that incorporated intervention in the variance equation and allowed for GARCH-in-mean. Neither effect was present. We did not include these specifications in the present paper because they greatly complicate estimation over the shorter subperiods.

5 Although our estimation period is January 6, 1983 to February 19, 1990, our data on intervention extend back to January 3, 1977.

consistent with the portfolio-balance approach, which predicts a significant relationship between relative asset stocks and the risk premium. In the portfolio-balance theory, the identity of countries intervening should not matter. Accordingly, our first measure of cumulative intervention, identified as CUM2 in the tables, combines the cumulative intervention of the two principal countries in each market, either the United States and Germany or the United States and Japan. The second portfolio-balance measure, CUMROW, is the cumulative aggregate intervention of the G10 countries less that of the two principal countries in each case. In our implementation of equation (3), we calculate the relative risk premium in terms of dollars. Negative (positive) values of the intervention data represent dollar sales (purchases), which, when sterilized, result in an increase (decrease) in dollar-denominated assets. Consequently, a negative coefficient on either CUM2 or CUMROW would be consistent with the portfolio-balance approach.[6]

In our tests of the impact of cumulative intervention, the null hypothesis is that markets are efficient and no portfolio-balance effect exists. Our use of cumulative, official intervention as a proxy for relative asset supplies implicitly assumes that investors have accurate information about intervention. If such is not the case, or if markets are inefficient, it is possible that intervention matters even if we do not reject our null. Thus, failure to reject leaves open the possibility that intervention is transmitting information to the market. As a further test of this possibility (and following Dominguez [1990]), we enter intervention without cumulating the data (intervention on date t-3 only). In this case, we also distinguish

6 See Edison (1990) and Weber (1986) for further discussions of this point.

between coordinated and unilateral intervention by separate countries. We define intervention as coordinated when the two principal central banks in a particular market undertake similar transactions on the same day. The tables that follow designate this type of intervention as COORD. We consider intervention to be unilateral if only one of the two principal countries transacts at a particular time, or if one buys while the other sells. In the tables, UNIL(A) refers to unilateral U.S. intervention, and UNIL(F) refers to unilateral German or Japanese intervention. Though we lack a theory implying particular signs for the coefficients on the coordinated and unilateral intervention terms, we expect stronger influences for the former, especially during periods when such agreements had been reached and widely publicized.

Data

Our sample period is January 6, 1983 to February 19, 1990, with 1,773 daily observations, excluding lags. We obtained exchange-rate and interest-rate data from the Paris market through DRIFACS PLUS (1988).[7] Yen-dollar and mark-dollar exchange rates are bid quotes as of 2:00 p.m. in Paris and are constructed as cross-rates for each currency quoted against the French franc. Interest rates, which we converted to a daily basis, are overnight Eurocurrency deposit rates, quoted on a 360-day basis, as of 9:30 a.m. The Paris market is the only one containing a complete set of data, most notably overnight Euroyen deposits.

---

[7] The ultimate source is Credit Lyonnais, Paris.

The Federal Reserve Board of Governors provided official intervention data. These include the net daily dollar transactions by Germany, Japan, and the United States, and the total for the G10 countries. Since the Board maintains all intervention transactions at their original dollar equivalents, we avoid any simultaneity problems associated with conversion. Over the period investigated, all U.S. intervention was against the mark or yen, except for a single purchase of $16.4 million equivalent British pounds in February 1985 (see Cross [1985, p. 58]).

We estimate the GARCH(1,1) model over the full sample period and over four subperiods that seem to represent different U.S. intervention regimes. During the first subperiod, January 6, 1983 to December 31, 1984, the United States rarely intervened, arguing that intervention was ineffectual. Germany, on the other hand, intervened often during this interval, while Japan intervened less frequently and undertook no unilateral transactions. The volume and frequency of U.S. intervention then increased markedly during the second subperiod, January 2, 1985 through December 31, 1985, especially following the Plaza Accord on September 23. The third subperiod, January 2, 1986 through February 20, 1987, marks a period of no U.S. intervention that ended with the Louvre Accord on February 23, 1987. Germany and Japan did intervene during this time. The volume and frequency of U.S., German, and Japanese intervention again rose sharply in the last subperiod, February 23, 1987 through February 19, 1990. A closer coordination of intervention also marked this last time frame.

Our empirical approach was to first estimate the basic structure of the model, exclusive of the intervention terms, over the full sample period.

We then maintained this basic model in all subsequent estimates, introducing intervention for the full sample period and for the four subperiods.

III. Results

The Basic Model

The first column of table 1 shows the results for the estimates of the basic model. Preliminary likelihood ratio tests supported 1) the inclusion of the MA(1) term, 2) the inclusion of each GARCH term (alpha, beta), and 3) the assumption of non-normality. Conditional non-normality is permitted by allowing the data to determine the value of the distributional term (1/v). The Ljung-Box Q statistics for the residuals, adjusted for heteroscedasticity ($Q/h_t{}^{.5}$), do not show significant serial correlation. The adjusted Q^2 statistics (Q^2/h_t) for the squared residuals further suggest that the GARCH specification is largely successful in removing heteroscedasticity. The sample statistics for skewness (B1) and kurtosis (B2), however, are larger than we would prefer. Under the null of conditional normality, B1 is distributed normally with mean zero and variance 6/n (n = number of observations). Under the same null, B2 is distributed normally with mean three and variance 24/n.

Unlike Dominguez (1990) and Loopesko (1984), we test for holiday and day-of-the-week effects in both the conditional mean and conditional variance equations, utilizing likelihood ratio tests to determine which dummies to retain for the subsequent estimations. The basic equations, to which we added intervention terms, are given in column 2 of table 1. Recall from above that we introduce intervention dated t-3 so that we can reasonably assume our risk premia (which involve variables dated t-2, t-1, and t) reflect intervention information available to market participants.

The dummy variables correspond to the t-1 date. For the mark-dollar risk premium, we find a Thursday effect in the mean and a U.S. holiday effect in the variance. The Thursday effect may reflect the tendency of U.S. and German monetary policy announcements to fall on this day. For the yen-dollar case, the conditional variance is higher on Fridays and on days following U.S. holidays. These results are similar to those of Baillie and Bollerslev (1989), who find a Friday effect in their analysis of exchange-rate changes.[8]

Intervention over the Full Sample Period

In column 3 of table 1, we introduce two portfolio-balance terms to the conditional-mean equation: the cumulative intervention of the principal central banks (denoted as CUM2) and cumulative intervention of the rest of the G10 (denoted as CUMROW). Because of possible collinearity between these terms, we examine their joint significance using likelihood ratio tests. For the mark-dollar case, the likelihood ratio (chi-squared) test with two degrees of freedom suggests that the cumulative intervention terms are jointly significant. For the yen-dollar case, the regressors are not jointly significant, suggesting no portfolio-balance effect in the yen-dollar market. We note, however, that both intervention coefficients are positive in the mark-dollar case. It is difficult to interpret this result as a portfolio-balance effect, since the logic of the portfolio-balance channel

---

8 See also So (1987) and McFarland, Petit, and Sung (1987).

implies that an increase in the supply of dollar assets raises the risk premium on dollars.

In column 4, we introduce coordinated and unilateral intervention flows. The results show that neither coordinated nor unilateral U.S. interventions are statistically significant in either market. Only unilateral German and Japanese interventions matter, and in each case, sales of dollars increase the risk premium.

Intervention over the Subperiods

For the 1983-84 subperiod, when the United States rarely intervened and undertook no unilateral transactions against the Japanese yen, likelihood ratio tests show that intervention was not statistically significant (see table 2). Although the results do not support the portfolio channel, we note that the volume of intervention was relatively small up to 1985. Within the context of the expectations channel, our findings suggest that central-bank intervention did not credibly communicate inside information during this period.

In 1985, intervention volumes increased sharply, particularly after September 23. During this time, a joint test of cumulative intervention terms is statistically significant in the mark-dollar case, but not in the yen-dollar case (see table 3). The individual coefficient on cumulative U.S. plus German intervention is statistically significant, but positive. When the model specification includes coordinated and unilateral interventions, only the former appears to be statistically significant for the yen-dollar case.

The United States did not intervene during 1986 (see table 4), and neither Japanese nor German cumulative interventions had a significant impact.

Moreover, when we redefine intervention as a flow, only unilateral Japanese intervention was statistically significant.

The volume and frequency of both U.S. and foreign intervention rose once again in our final subperiod as countries attempted to coordinate interventions more closely (see table 5). Here, the cumulative intervention terms in the mark-dollar case are jointly significant, but only at a 10 percent confidence level. In the yen-dollar case, the cumulative intervention terms are not significant. When we redefine intervention as coordinated and unilateral flows, only unilateral German intervention is statistically significant.

IV. Conclusion

Using official data on daily dollar intervention and a GARCH specification to model the deviation from UIP, we test for the impact of intervention on foreign exchange risk premia. We adopt two general specifications of intervention. In one, we calculate cumulative intervention, which is broadly consistent with the portfolio-balance approach. In the other, we consider coordinated and unilateral intervention flows, which are broadly consistent with the expectations channel. By examining subperiods, we hope to control for shifts in intervention policies that could confound results for the full sample period.

As with other empirical studies of intervention, the results are mixed. None of the intervention terms is consistently significant across all time periods, and we find little support for the portfolio-balance channel. Dollar intervention took place in each subperiod, and from the perspective of the portfolio-balance approach, we would expect the coefficients on the cumulative intervention terms to be consistently significant. Moreover, these coefficients were typically positive.

When we examine the expectations channel, we reach the surprising conclusion that unilateral intervention is significant more often than coordinated intervention. One interpretation of this would be that coordination did not clarify exchange-rate policies or intentions. An exception to this interpretation was the coordination of U.S. and Japanese intervention in 1985, the year of the Plaza Accord. At that time, the market expected fundamental changes in G3 monetary and fiscal policy, but, as Humpage (1988) indicates, Germany seemed less willing than Japan to follow through. We also find that unilateral U.S. intervention is never significant and that in the 1983-84 subperiod, when the United States discredited an active intervention policy, no intervention variable seemed significant.

References

Backus, D., and Kehoe, P., "On the Denomination of Government Debt: A Critique of the Portfolio Balance Approach," Federal Reserve Bank of Minneapolis, Working Paper No. 359, 1988.

Baillie, Richard T., and Bollerslev, Tim, "The Message in Daily Exchange Rates: A Conditional-Variance Tale," Journal of Business and Economic Statistics, 7 (July 1989): 297-305.

Baillie, Richard T., and McMahon, Patrick C., The Foreign Exchange Market: Theory and Econometric Evidence. Cambridge, England: Cambridge University Press, 1989.

Baillie, Richard T., and Osterberg, William P., "The Risk Premium in Forward Foreign Exchange Markets and G-3 Central Bank Intervention: Evidence of Daily Effects, 1985-1990," Federal Reserve Bank of Cleveland, Working Paper No. 9109, July 1991.

Cross, Sam Y., "Treasury and Federal Reserve Foreign Exchange Operations, February-July 1985 Semiannual Report," Federal Reserve Bank of New York, Quarterly Review, Autumn 1985: 50-64.

Cumby, Robert E., "Is It Risk? Explaining Deviations from Uncovered Interest Parity," Journal of Monetary Economics, 22 (September 1988): 279-299.

Dominguez, K., "The Informational Role of Official Foreign Exchange Intervention Operations," manuscript, Kennedy School of Government, Harvard University, 1988.

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

________, and Frankel, Jeffrey, "Does Foreign Exchange Intervention Matter? Disentangling the Portfolio and Expectations Effects for the Mark," National Bureau of Economic Research Working Paper No. 3299, 1990.

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

Edison, Hali J., "Foreign Currency Operations: An Annotated Bibliography," Board of Governors of the Federal Reserve System, International Finance Discussion Papers 380, May 1990.

Engel, Charles M., and Rodrigues, Anthony, P., "Tests of International CAPM with Time-Varying Covariances," National Bureau of Economic Research Working Paper No. 2303, July 1987.

Frankel, Jeffrey A., and Froot, Kenneth, "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations," American Economic Review, 7 (March 1987): 133-153.

Ghosh, Atish R., "Is It Signalling? Exchange Intervention and the Dollar-Deutschemark Rate," Princeton University, Working Paper, 1989.

Giovannini, Alberto, and Jorion, Philippe, "The Time Variation of Risk and Return in the Foreign Exchange and Stock Markets," Journal of Finance, 44 (June 1989): 307-325.

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

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

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

________, "Modeling Heteroscedasticity in Daily Foreign-Exchange Rates," Journal of Business and Economic Statistics, 7 (July 1989): 307-317.

Humpage, Owen F., "Intervention and the Dollar's Decline," Federal Reserve Bank of Cleveland, Economic Review, 24 (1988 Quarter 2): 2-16.

Hung, Juann H., "Noise Trading and the Effectiveness of Sterilized Foreign Exchange Intervention," Federal Reserve Bank of New York, Research Paper No. 9111, March 1991a.

________, "The Effectiveness of Sterilized U.S. Foreign Exchange Intervention -- An Empirical Study Based on the Noise Trading Approach," Federal Reserve Bank of New York, Working Paper No. 9118, May 1991b.

Ito, Takatoshi, "Foreign Exchange Rate Expectations: Micro Survey Data," American Economic Review, 80 (June 1990): 434-449.

Klein, Michael, and Rosengren, Eric, "What Do We Learn from Foreign Exchange Intervention?" Federal Reserve Bank of Boston, Working Paper, September 1991.

Lastrapes, William D., "Exchange Rate Volatility and U.S. Monetary Policy: An ARCH Application," Journal of Money, Credit, and Banking, 21 (February 1989): 66-77.

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

Lucas, Robert E., Jr., "Interest Rates and Currency Prices in a Two-Country World," Journal of Monetary Economics, 10 (November 1982): 335-360.

Lyons, Richard K., "Tests of the Foreign Exchange Risk Premium Using the Expected Second Moments Implied by Option Pricing," Journal of International Money and Finance, 7 (March 1988): 91-108.

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

McCurdy, T.H., and Morgan, I.G., "Testing the Martingale Hypothesis in Deutschemark Futures with Models Specifying the Form of Heteroscedasticity," Journal of Applied Econometrics, 3 (July-September 1988): 187-202.

McFarland, J.W., Petit, R.R., and Sung, S.K., "The Distribution of Foreign Exchange Price Changes: Trading Day Effects and Risk Measurement: A Reply," Journal of Finance, 42 (1987): 189-194.

Nelson, D.B., "Conditional Heteroscedasticity and Asset Returns: A New Approach," mimeo, Massachusetts Institute of Technology, 1987.

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

Osterberg, William, "Intervention and the Risk Premium in Foreign Exchange Rates," Federal Reserve Bank of Cleveland, Working Paper No. 8908, August 1989.

Pagan A., and Hong, Y., "Non-parametric Estimation and the Risk Premium," University of Rochester Center for Economic Research, Working Paper No. 135, 1988.

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

So, J.C., "The Distribution of Foreign Exchange Price Changes: Trading Day Effects and Risk Measurements -- A Comment," Journal of Finance, 42 (1987): 181-188.

Weber, Warren E., "Do Sterilized Interventions Affect Exchange Rates?" Federal Reserve Bank of Minneapolis, Quarterly Review, Summer 1986: 14-23.

TABLE 1: FULL-SAMPLE ESTIMATES OF BASIC MODEL WITH INTERVENTION

(n = 1773)

	MARK-DOLLAR EXCHANGE RATE				YEN-DOLLAR EXCHANGE RATE			
	1	2	3	4	1	2	3	4
I. CONDITIONAL MEAN								
constant	0.00012	0.00033	0.00179	0.00022	0.00016	0.00012	0.00052	0.00012
	(0.75135)	(2.00822)	(4.20291)	(1.31652)	(1.44011)	(1.08310)	(0.97004)	(1.11198)
THURSDAY		-1.1E-03	-1.1E-03	-9.9E-04				
		(-2.66826)	(-2.66049)	(-2.45181)				
CUM2			5.7E-04				-6.0E-05	
			(3.98169)				(-0.48548)	
CUMROW			6.3E-05				1.1E-04	
			(1.61586)				(0.77177)	
COORD				-3.2E-03				-7.5E-04
				(-0.21868)				(-1.11801)
UNIL(A)				3.3E-02				6.8E-02
				(0.28008)				(0.71710)
UNIL(F)				-1.1E-01				-4.7E-03
				(-5.05312)				(-5.23908)
MA(1)	-0.05230	-0.05189	-0.06329	-0.05725	-0.09772	-0.10200	-0.10270	-0.10848
	(-2.14581)	(-2.11541)	(-2.57961)	(-2.31253)	(-4.06845)	(-4.16959)	(-4.17953)	(-4.42164)
II. CONDITIONAL VARIANCE								
omega	2.5E-06	2.1E-06	2.3E-06	2.2E-06	3.9E-06	1.6E-06	1.7E-06	1.9E-06
	(2.96386)	(2.64310)	(2.72888)	(2.69636)	(4.48789)	(1.36865)	(1.42442)	(1.51865)
alpha	0.08570	0.09145	0.09442	0.09354	0.20582	0.22182	0.22226	0.22969
	(4.77381)	(4.93600)	(5.00707)	(5.01063)	(5.40014)	(5.57873)	(5.59132)	(5.51546)
beta	0.87768	0.86484	0.85869	0.86021	0.75667	0.70490	0.70330	0.68899
	(34.54128)	(32.95260)	(31.80755)	(32.35567)	(24.45744)	(19.07988)	(18.94772)	(17.82338)
HOLIDAY		2.1E-05	2.1E-05	2.1E-05		3.0E-05	3.0E-05	3.3E-05
		(3.04623)	(3.08038)	(3.13940)		(2.98241)	(2.97558)	(3.02876)
FRIDAY						1.4E-05	1.3E-05	1.4E-05
						(2.96671)	(2.92760)	(3.13211)
1/v	0.17742	0.17122	0.16416	0.16151	0.27972	0.26334	0.26276	0.26701
	(16.03262)	(14.96946)	(6.12272)	(8.40552)	(404.95671)	(437.46791)	(12.23973)	(9.69089)
III. DIAGNOSTICS								
log likelihood	6201.58013	6212.83696	6220.74443	6220.69995	6544.03015	6557.11599	6557.54803	6562.11156
B1 for E/SQRT(H)	-0.36182	-0.47712	-0.43288	-0.49028	-0.86635	-0.78347	-0.78970	-0.78076
B2 for E/SQRT(H)	6.31236	6.52908	6.36310	6.48799	22.80177	22.59273	22.58761	21.98630
Q(15) for E	17.77931	17.35026	15.11327	14.85418	22.63084	22.26044	21.48530	19.29524
Q(15) for E/SQRT(H)	19.94990	21.63347	15.71453	14.11179	21.55567	20.80021	19.94822	16.82556
Q(15) for E**2	57.21022	55.77838	58.51609	56.34464	294.00307	295.65808	296.71499	299.91613
Q(15) for E**2/(H)	15.29835	19.00623	18.60139	20.62425	2.86040	2.20594	2.21587	2.14139

CUM2: U.S. plus German or U.S. plus Japanese intervention cumulated from January 2, 1977.

CUMROW: Group of Ten countries plus Switzerland less relevant CUM2 countries cumulated from January 2, 1977.

COORD: Coordinated U.S. and German or U.S. and Japanese intervention.

UNIL(A): Unilateral U.S. intervention.

UNIL(F): Unilateral German or unilateral Japanese intervention.

HOLIDAY: Dummy variable for U.S. holidays at t-1.

THURSDAY: Thursday dummy variable.

FRIDAY: Friday dummy variable.

Note: t-statistics are in parentheses.

Data Sources: DRIFACS PLUS; and BOARD OF GOVERNORS OF THE FEDERAL RESERVE SYSTEM.

TABLE 2: BASIC MODEL AND INTERVENTION, JANUARY 6, 1983 TO DECEMBER 31, 1984

(n = 496)

	1983 -- 1984					
	MARK-DOLLAR			YEN-DOLLAR		
	1	2	3	1	2	3
I. CONDITIONAL MEAN						
constant	0.00127	0.01011	0.00112	0.00037	-0.00129	0.00037
	(4.42800)	(1.78993)	(3.55144)	(2.36137)	(-0.62635)	(2.34026)
THURSDAY	-1.6E-03	-1.6E-03	-1.6E-03			
	(-2.12495)	(-2.15904)	(-2.12364)			
CUM2		-1.8E-05			-5.5E-03	
		(-0.06531)			(-1.20976)	
CUMROW		2.3E-03			1.0E-03	
		(1.57211)			(0.82828)	
COORD			5.4E-03			-2.1E-03
			(0.04848)			(-0.05048)
UNIL(A)			-9.4E-02			
			(-0.02839)			
UNIL(F)			-5.6E-02			-2.0E-05
			(-1.28699)			(-0.00091)
MA(1)	-0.11567	-0.11900	-0.11499	-0.16305	-0.16744	-0.16314
	(-2.37951)	(-2.42716)	(-2.34892)	(-3.57769)	(-3.69845)	(-3.5769)
II. CONDITIONAL VARIANCE						
omega	3.1E-06	4.8E-06	3.5E-06	-1.8E-08	1.0E-09	-2.0E-08
	(1.45711)	(1.57798)	(1.46510)	(-0.01231)	(0.00100)	(-0.01376)
alpha	0.11470	0.13064	0.11914	0.10349	0.10789	0.10369
	(2.81371)	(2.81606)	(2.79562)	(2.28635)	(2.31995)	(2.28231)
beta	0.81635	0.76344	0.80341	0.83563	0.83674	0.83573
	(12.74161)	(8.99443)	(11.37179)	(12.85556)	(13.15144)	(12.84379)
HOLIDAY	1.6E-05	2.1E-05	1.8E-05	8.0E-06	8.5E-06	8.0E-06
	(1.59428)	(1.74612)	(1.66045)	(1.33194)	(1.38076)	(1.33200)
FRIDAY				7.1E-06	6.6E-06	7.1E-06
				(1.32508)	(1.21553)	(1.32139)
1/v	0.14267	0.14265	0.14341	0.20645	0.21569	0.20680
	(13.51146)	(3.29889)	(7.44277)	(137.07427)	(121.19414)	(5.96222)
III. DIAGNOSTICS						
log likelihood	1777.42770	1778.93421	1778.42566	1953.20183	1954.32670	1953.13615
B1 for E/SQRT(H)	-0.43643	-0.47107	-0.44120	-2.50129	-2.63276	-2.50404
B2 for E/SQRT(H)	4.80685	4.82936	4.79891	28.28391	30.18254	28.33324
Q(15) for E	16.73905	16.71905	16.94027	10.63930	10.66125	10.61000
Q(15) for E/SQRT(H)	15.28873	15.40371	14.98315	9.13170	9.01208	9.06887
Q(15) for E**2	21.12632	20.59913	19.32861	7.67696	7.73975	7.67776
Q(15) for E**2/(H)	11.15610	11.99264	11.48663	0.68292	0.66203	0.68215

KEY: See table 1.

Note: t-statistics are in parentheses.

Data Sources: DRIFACS PLUS; and BOARD OF GOVERNORS OF THE FEDERAL RESERVE SYSTEM.

TABLE 3: BASIC MODEL AND INTERVENTION, JANUARY 2, 1985 TO DECEMBER 31, 1985
(n = 247)

	1985					
	MARK-DOLLAR			YEN-DOLLAR		
	1	2	3	1	2	3
I. CONDITIONAL MEAN						
constant	-0.00033	0.01182	-0.00061	-0.00015	0.00678	-0.00003
	(-0.55410)	(1.92813)	(-1.1491)	(-0.53884)	(1.61586)	(-0.1261)
THURSDAY	-2.6E-04	-4.0E-04	-1.0E-04			
	(-0.19661)	(-0.31161)	(-0.08633)			
CUM2		5.1E-03			1.2E-03	
		(2.30424)			(0.71417)	
CUMROW		-6.7E-04			1.1E-03	
		(-0.40243)			(1.18690)	
COORD			-3.0E-02			1.5E-02
			(-0.78501)			(3.0084)
UNIL(A)			-1.7E-01			-2.8E-03
			(-0.46572)			(-0.00864)
UNIL(F)			-3.4E-01			5.0E-03
			(-1.39325)			(0.43606)
MA(1)	0.02153	0.00851	0.00524	-0.03987	-0.04744	-0.04161
	(0.31918)	(0.12501)	(0.07762)	(-0.60921)	(-0.7134)	(-0.63447)
II. CONDITIONAL VARIANCE						
omega	5.4E-06	4.5E-06	5.0E-06	2.5E-06	2.4E-06	1.8E-06
	(1.15776)	(1.13956)	(1.14305)	(0.71100)	(0.70080)	(0.53907)
alpha	0.13205	0.12812	0.14496	0.17402	0.17523	0.15912
	(1.82704)	(1.83997)	(1.95133)	(1.69266)	(1.67985)	(1.55917)
beta	0.81843	0.83296	0.81130	0.66030	0.65619	0.69221
	(9.0524)	(10.06786)	(9.34928)	(4.69268)	(4.61023)	(5.11104)
HOLIDAY	2.0E-05	2.2E-05	3.0E-05	3.3E-05	3.2E-05	2.9E-05
	(0.57471)	(0.56432)	(0.71044)	(1.45337)	(1.43366)	(1.34633)
FRIDAY				1.5E-05	1.6E-05	1.5E-05
				(1.39427)	(1.42762)	(1.40228)
1/v	0.24297	0.25871	0.25373	0.26419	0.26211	0.26498
	(5.21579)	(4.19618)	(8.2578)	(40.29954)	(5.36840)	(7.38294)
III. DIAGNOSTICS						
log likelihood	814.59870	818.18390	817.35638	941.25145	943.21163	944.39614
B1 for E/SQRT(H)	-1.40943	-1.48471	-1.38631	-4.18447	-4.23100	-4.27508
B2 for E/SQRT(H)	11.39616	11.95683	11.64777	42.94233	43.79033	44.30965
Q(15) for E	17.27667	17.97093	16.57776	23.17124	24.09430	21.13707
Q(15) for E/SQRT(H)	13.67557	14.46309	13.66817	11.03675	11.34871	12.05467
Q(15) for E**2	6.58222	6.32270	7.10894	50.95694	50.72152	11.03209
Q(15) for E**2/(H)	6.62992	5.94628	6.74183	1.19978	1.10907	0.58586

KEY: See table 1.

Note: t-statistics are in parentheses.

Data Sources: DRIFACS PLUS; and BOARD OF GOVERNORS OF THE FEDERAL RESERVE SYSTEM.

TABLE 4: BASIC MODEL AND INTERVENTION, JANUARY 2, 1986 TO FEBRUARY 20, 1987

(n = 285)

	1986-1987					
	MARK-DOLLAR			YEN-DOLLAR		
	1	2	3	1	2	3
I. CONDITIONAL MEAN						
constant	-0.00089	0.00902	-0.00086	-0.00041	0.00255	-0.00027
	(-1.68086)	(0.42062)	(-1.60133)	(-1.32947)	(0.26754)	(-1.04524)
THURSDAY	-1.1E-03	-1.1E-03	-9.6E-04			
	(-0.74966)	(-0.76248)	(-0.66376)			
CUM2		3.5E-03			8.2E-04	
		(0.43241)			(1.02226)	
CUMROW		-1.6E-04			4.1E-04	
		(-0.16758)			(0.22220)	
COORD						
UNIL(A)						
UNIL(F)			-3.3E-02			-5.1E-03
			(-0.25442)			(-5.62539)
MA(1)	-0.06771	-0.06269	-0.06928	-0.13803	-0.14963	-0.15965
	(-1.08541)	(-1.03440)	(-1.10344)	(-2.30780)	(-2.47785)	(-2.78923)
II. CONDITIONAL VARIANCE						
omega	6.7E-06	6.4E-06	6.8E-06	3.9E-06	-1.4E-06	1.4E-06
	(1.11916)	(1.05993)	(1.10699)	(0.61004)	(-0.28008)	(0.23821)
alpha	0.07831	0.07347	0.07720	0.42300	0.44878	0.54855
	(1.45019)	(1.36299)	(1.42306)	(2.02319)	(1.98414)	(1.69373)
beta	0.82004	0.82630	0.81959	0.48600	0.52536	0.48574
	(7.23083)	(7.07413)	(7.04926)	(4.40443)	(5.49557)	(5.07926)
HOLIDAY	5.1E-05	5.1E-05	5.0E-05	1.2E-04	1.4E-04	1.7E-04
	(1.29967)	(1.28442)	(1.28323)	(0.83603)	(0.92834)	(0.87586)
FRIDAY				9.6E-05	1.2E-04	1.1E-04
				(2.32708)	(2.26765)	(1.84733)
1/v	0.06104	0.09441	0.05957	0.36154	0.34205	0.36992
	(27.97599)	(8.48291)	(7.26717)	(5398.35104)	(22.89131)	(287.17127)
III. DIAGNOSTICS						
log likelihood	949.19250	949.10856	949.29937	975.82996	978.12679	981.71062
B1 for E/SQRT(H)	0.06976	0.05904	0.06267	1.96757	2.04057	1.99512
B2 for E/SQRT(H)	3.44826	3.44142	3.43644	17.60197	17.66666	16.56339
Q(15) for E	14.69948	14.69844	14.76686	15.63504	15.39223	16.19823
Q(15) for E/SQRT(H)	15.41799	15.33652	15.43139	12.22355	9.90027	11.94551
Q(15) for E**2	24.15303	23.55600	23.89456	67.14251	67.94963	68.91197
Q(15) for E**2/(H)	13.72062	13.36345	13.75257	2.56712	3.31988	6.43461

KEY: See table 1.

Note: t-statistics are in parentheses.

Data Sources: DRIFACS PLUS; and BOARD OF GOVERNORS OF THE FEDERAL RESERVE SYSTEM.

TABLE 5: BASIC MODEL AND INTERVENTION, FEBRUARY 23, 1987 TO FEBRUARY 19, 1990

(n = 745)

	1987-1990					
	MARK-DOLLAR			YEN-DOLLAR		
	1	2	3	1	2	3
I. CONDITIONAL MEAN						
constant	0.00010	0.00022	0.00010	0.00009	-0.00166	0.00012
	(0.46682)	(0.26310)	(0.46102)	(0.58313)	(-1.35552)	(0.60178)
THURSDAY	-8.3E-04	-8.1E-04	-8.1E-04			
	(-1.50040)	(-1.50464)	(-1.46713)			
CUM2		2.4E-04			4.5E-04	
		(1.28279)			(1.53338)	
CUMROW		1.5E-04			-2.8E-04	
		(1.58822)			(-1.15004)	
COORD			-2.4E-05			-1.1E-03
			(-0.00341)			(-1.58236)
UNIL(A)			5.8E-02			1.3E-01
			(0.47833)			(1.25258)
UNIL(F)			-1.3E-01			-1.9E-03
			(-4.20748)			(-0.73284)
MA(1)	-0.06319	-0.07069	-0.06846	-0.08443	-0.09200	-0.08570
	(-1.68892)	(-1.88764)	(-1.80276)	(-2.25208)	(-2.34031)	(-2.16197)
II. CONDITIONAL VARIANCE						
omega	1.7E-06	1.8E-06	1.7E-06	2.1E-06	2.6E-06	3.1E-06
	(1.74063)	(1.75793)	(1.75889)	(0.87074)	(1.08793)	(1.29356)
alpha	0.05573	0.05648	0.05503	0.17044	0.18543	0.18889
	(2.57286)	(2.55935)	(2.50243)	(3.28812)	(3.43146)	(3.40594)
beta	0.89218	0.88986	0.89085	0.74783	0.71934	0.70567
	(24.80069)	(23.91686)	(24.19857)	(11.57685)	(10.56917)	(10.10768)
HOLIDAY	2.3E-05	2.3E-05	2.3E-05	3.3E-05	3.5E-05	3.8E-05
	(2.48388)	(2.46179)	(2.49657)	(2.04474)	(2.11614)	(2.14088)
FRIDAY				1.1E-05	1.1E-05	1.0E-05
				(1.31476)	(1.41686)	(1.31914)
1/v	0.17317	0.16992	0.15969	0.24593	0.22515	0.22152
	(9.38930)	(6.89037)	(4.56771)	(5.35029)	(5.09048)	(25.52472)
III. DIAGNOSTICS						
log likelihood	2692.50881	2694.97306	2695.71505	2708.31798	2709.83988	2710.35265
B1 for E/SQRT(H)	-0.27914	-0.23575	-0.32264	-0.24324	-0.21885	-0.26799
B2 for E/SQRT(H)	5.69759	5.47859	5.54766	5.61619	5.52321	5.80918
Q(15) for E	17.04329	14.65995	13.32945	21.86380	21.54048	20.56957
Q(15) for E/SQRT(H)	14.89185	11.36793	10.08280	25.46855	23.72444	21.70662
Q(15) for E**2	37.34961	37.83676	36.99809	89.05132	89.30957	86.57824
Q(15) for E**2/(H)	21.00178	20.51776	20.49623	13.73176	13.97497	13.79804

KEY: See table 1.

Note: t-statistics are in parentheses.

Data Sources: DRIFACS PLUS; and BOARD OF GOVERNORS OF THE FEDERAL RESERVE SYSTEM.