New Results on the Rationality of Survey Measures of Exchange-Rate Expectations

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Introduction

Analyses of economic and financial developments often rely on propositions about the rationality of market participants. Particularly in financial markets, where information is widely and readily disseminated, it is commonly presumed that economic agents have "rational expectations" about the future course of events. This assumption can have powerful implications for the efficacy of certain government policies.

Questioning the rationality of foreign-exchange market participants is particularly tempting, in light of the widely acknowledged poor performance of many economists' models of exchange-rate determination. Models that view exchange rates as equal to the expected present discounted value of future "fundamentals" (for instance, monetary policy, fiscal policy, and trade flows) and assume that participants have rational expectations about future fundamentals have done poorly in predicting exchange-rate movements. Although this latter finding has stimulated a wide body of research, the validity of the *rational expectations hypothesis* remains unresolved.

An anomaly in international finance related to the rational expectations hypothesis is the forward discount puzzle, in which the forward foreign exchange rate predicts the wrong direction of movement for the future spot rate. Most verifications of this puzzle presume rational expectations and, thus, are worth reconsidering if the hypothesis is to be rejected. Lewis (1995) summarizes work surrounding this puzzle. Baillie and Bollerslev (1997) examine earlier findings, suggesting that with more recent data the puzzle no longer appears.

Dominguez and Frankel (1993) illustrate how crucial the rational expectations assumption might be in analyzing the impact of government policy. They conclude that central banks' foreign-exchange intervention had a significant impact on risk premia in currency markets during 1982–88. However, their finding hinges on survey data on exchange-rate expectations. When rational expectations are imposed, intervention is seen to be ineffective.

In this article, I extend previous research on the rationality of survey measures of expectations for foreign exchange rates. Section I reviews the literature, highlighting several interpretations of rationality. Section II provides some motivation for my choice of econometric tests. Section III presents summary information about the data and then the results of the main tests of interest. Finally, I summarize what has been gained from the exercise and what might be suggested for future research.

I. Related Literature

Econometric analyses of survey data have typically focused on the rationality of participants' expectations. The most familiar interpretation of rationality is expressed in terms of expectations representing unbiased forecasts of the actual future outcome. However, several related questions might be of interest. For example, one might question whether expectations incorporate—or react to—news of fundamentals. In addition, rationality might imply a specific relationship between short-run and long-run expectations. These questions suggest the value of studying alternative mechanisms through which expectations are formed, and they are closely related to the issue of whether a risk premium exists in foreign exchange markets.[1]

In his survey of surveys, Takagi (1991) notes three characteristics of survey data on expectations of future exchange rates. First, the dispersion of expectations tends to increase with the forecast horizon, an outcome that may be related to group effects.[2] Interestingly, Ito (1993) finds that Japanese exporters had expectations of greater yen depreciation, while Japanese importers had exactly the opposite expectations.[3] Second, expected changes in exchange rates tend to underpredict the actual extent of exchange-rate movements, implying that much of actual exchange-rate movements are unexpected. The third characteristic is referred to as "twist"—that is, longer-run expectations tend to reverse the direction of short-run expectations. For example, an appreciation would tend to be followed by an expectation of further depreciation, but an expectation of further appreciation in the more distant future.

The "Unbiasedness" Interpretation of Rationality

This paper will focus on the most familiar interpretation of rationality—that survey measures are unbiased forecasts of actual future outcomes. Dominguez (1986) tested this hypothesis by regressing actual depreciation on expected depreciation for the 1983–85 period using data from both Money Market Services and the Japanese Center for International Finance. Her results strongly reject the "unbiasedness" hypothesis with one-week, one-month, and three-month data. Ito (1993) also rejected the hypothesis for the 1985–87 period, at least for longer-run horizons. Cavaglia, Verschoor, and Wolff (1993) confirm this finding with EMS exchange rates, as well as with exchange rates against the U.S. dollar for 1986–90.[4] Beng and Siong (1993) also reject unbiasedness for the Singapore currency against the dollar for 1984–91 for all forecast horizons. However, Liu and Maddala (1992), using cointegration techniques, cannot reject the rational expectations hypothesis for the 1984–89 period.

Orthogonality

The second most familiar interpretation of the rational expectations hypothesis is that expectations incorporate all available information. Takagi (1991) summarizes examinations of this "orthogonality" hypothesis. Generally, Dominguez (1986), Froot and Frankel (1989), Ito (1993), MacDonald and Torrance (1989), Cavaglia, Verschoor, and Wolff (1993), and Beng and Siong (1993) find that the survey data do not fully incorporate all available information.

Long Run versus Short Run

Findings of a twist (short-run expectations show bandwagon effects, but long-run expectations are stabilizing) motivate an examination of the connection between the short run and long run. Froot and Ito (1989) propose a defini-

■ **1** The use of survey data to extract risk premia or to study the mechanisms through which expectations might be formed will not be discussed here, except to note that several of the articles surveyed discuss the issues. Among those that use survey data to extract risk premia are Frankel and Froot (1987a) and Dutt and Ghosh (1995). Among studies of mechanisms, see Frankel and Froot (1987a, 1987b).

■ **2** Cavaglia, Verschoor, and Wolff (1993) find that the mean expected depreciation tends to fall with the forecast horizon, as does the variance of forecast errors. This is contrary to the findings of other researchers.

■ **3** Ito (1993) is the only study to use panel data and to examine heterogeneity among survey respondents. He also finds individual as well as industry effects. As Ito points out, if all relevant information about exchange rates is public, the finding of heterogeneity implies a rejection of rationality.

■ **4** These authors remind us that the "peso problem"—wherein market participants allow for a small probability of a large change in the future exchange rate—can explain an ex post finding of bias even if expectations are formed rationally.

tion of consistency. However, any definition of consistency requires the mechanism through which expectations are formed to be specified. Their evidence is mixed with results that differ with horizon. Ito (1993) confirms the existence of twist and also rejects consistency. However, in their study of Singapore's currency, Beng and Siong (1993) find no evidence of twist.

Expectations Formation

Takagi (1991) summarizes studies of the mechanisms of expectations formation by Frankel and Froot (1987a,b), the Bank of Japan (1989), and Froot and Frankel (1990). *Extrapolative expectations* means that the expected currency movement is related to the most recent movement. Generally, examination of this mechanism leads to the conclusion that bandwagon effects are present in the short run, but effects of the opposite sign are present for longer horizons. The effects are usually stabilizing. With *adaptive expectations*, the expected movement represents an average of the actual current and the expected current. The results of Frankel and Froot (1987a, b) are inconclusive regarding the validity of this mechanism. The length of the horizon seems to matter, and findings are not inconsistent with an unanticipated appreciation leading to an expected depreciation in the long run. *Regressive expectations* are said to exist when the actual exchange rate is expected to move toward an equilibrium rate. The results of Frankel and Froot (1987a, b), the Bank of Japan (1989), and Froot and Frankel (1990) point to the conclusion that expectations can be destabilizing in the short run, moving away from equilibrium, whereas the opposite effect occurs for longer horizons. Beng and Siong (1993) examine the same three mechanisms and find stabilizing extrapolative expectations but no stabilizing adaptive mechanism operative; in terms of the regressive mechanism, both short-run and long-run expectations move backward toward an equilibrium value. Ito (1994) finds that, despite the presence of mean reversion in the actual exchange-rate series for the yen, such reversion was not captured in a six-month horizon for expectations.

Chartists and Fundamentalists

Froot and Frankel (1990) and Allen and Taylor (1990) have suggested that the differences between short-run and long-run expectations might be related to the types of forecasting techniques employed. In particular, it is reasonable to speculate that short-run forecasts are derived from "chartists," or technical analysis, while longer-run forecasts are based on models of fundamentals. This possibility has been largely unexplored (as of Takagi [1991]). Hung (1997) relates the activities of chartists to the conduct of U.S. central bank foreign-exchange intervention policy. Bhattacharya and Weller (1997) explore possible implications of asymmetric information for intervention.

II. Test Specifications

In this article I test the unbiasedness hypothesis using Money Market Services data for 1989–97 on the deutsche mark/dollar (DM/$) exchange rate. My results will update Liu and Maddala's (1992) analysis of 1984–89. In order to properly compare our results with those previously published, I will review the progression of econometric techniques that have been utilized in this area.

At first one might be tempted to estimate the equation

$$(1) \qquad S_{t+k} = \alpha + \beta S^{e}_{t,t+k} + e_{t+k},$$

where S_{t+k} is the actual future exchange rate and $S^{e}_{t,t+k}$ is the expectation at time t of the future exchange rate at $t+k$.[5] The unbiasedness hypothesis stipulates that $\alpha = 0$ and $\beta = 1$. In addition, we might look at the orthogonality condition, regressing $(S_{t+k} - S^{e}_{t,t+k})$ on information available at t and testing for coefficients equal to zero. Or, we might test to see if forecast errors were serially uncorrelated. If the latter is not the case, it would imply that the forecast could be improved by considering past errors. However, as a wide body of research has discussed, if S_{t+k} and $S^{e}_{t,t+k}$ are nonstationary and follow unit-root processes, conventional t tests will be incorrect. To avoid this, some have suggested estimating equation (2)

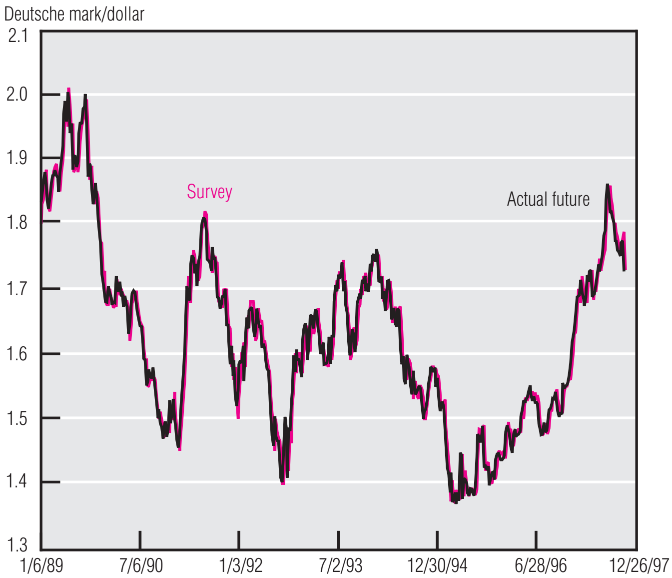
$$(2) \qquad S_{t+k} - S_t = \alpha + \beta (S^{e}_{t,t+k} - S_t) + e_{t+k},$$

and testing $\alpha = 0$ and $\beta = 1$. Although the left side of equation (2) is stationary, if both components have unit roots it is not clear that the same can be said for the right side. By analogy

■ 5 For the time being, I ignore the serial correlation issues that arise when k is not equal to one when the sampling frequency is one.

FIGURE 1

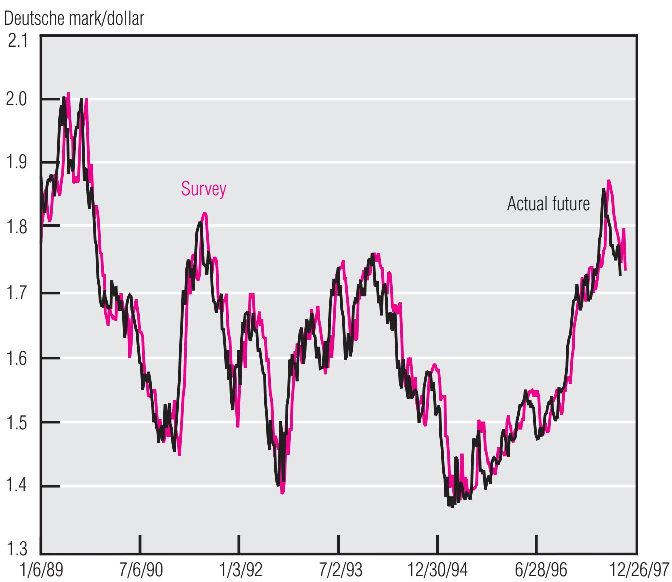
One-Week-Ahead Exchange Rate



SOURCE: Author's calculations.

FIGURE 2

One-Month-Ahead Exchange Rate



SOURCE: Author's calculations.

to the arguments of Liu and Maddala, the right side can be written as $(S^e_{t,t+k} - S^e_{t-1,t+k-1}) + (S^e_{t-1,t+k-1} - S_t)$; the second term is stationary only if $\alpha = 0$ and $\beta = 1$. Another possibility is to estimate equation (3),

$$(3) \quad S_{t+k} - S_{t+k-1} = \alpha + \beta(S^e_{t,t+k} - S^e_{t-1,t+k-1}) + e_{t+k},$$

which omits $\upsilon(S^e_{t-1,t+k-1} - S_t)$ from the right side.

I will follow the more direct approach suggested by Liu and Maddala. First, note that if S_{t+k} follows a random walk, so should its rational forecast; thus, the two time series should be cointegrated with a factor of 1 and random residuals. However, standard cointegration tests allow estimation of the cointegrating factor and do not require random-error terms. Rather than estimate equation (1) as the cointegrating equation, we first test whether S_{t+k} and $S^e_{t,t+k}$ are unit-root processes, and then test for the stationarity of $\mu_t = (S_{t+k} - S^e_{t,t+k})$. The second step restricts $\alpha = 0$ and $\beta = 1$, so that it can be referred to as a restricted cointegration test. As Liu and Maddala state, if μ_t is stationary, then S_{t+k} and $S^e_{t,t+k}$ are cointegrated with a factor of 1 because the cointegrating factor is unique when it exists.

Testing for Unit Roots

The most widely used unit-root tests for a variable y_t rely on equations of the form

$$(4) \quad \Delta y_t = \delta_0 + \delta\, t + (\alpha - 1)y_{t-1} + \sum_{j=1}^{p} \alpha_j \Delta y_{t-j} + e_t$$

and focus on the coefficient associated with y_{t-1}. A key consideration is how many lags of the left-side variable to include. Liu and Maddala follow a procedure suggested by Schwert (1989). Ng and Perron (1995) describe two popular procedures, the Akaike information criterion and a criterion suggested by Schwartz, both of which minimize an objective function of the form

$$(5) \quad I_k = \log\, \hat{\sigma}^2_k + k\, C_T / T,$$

where σ^2 is the maximum-likelihood estimate of the variance, T is the number of observations, and k is the number of right-side variables. For the Akaike information criterion, $C_T = 2$, and for the Schwartz criterion, $C_T = \log(T)$. Ng and

TABLE 1

Unit-Root Tests on Logarithms of Future and Expected Future Deutsche Mark/Dollar Exchange Rates

Statistics	Weekly		Monthly	
	Actual	Survey	Actual	Survey
τ^{τ} (LR)[a]	–2.429	–2.117	–2.430	–2.288
τ^{τ} (BIC)[b]	–1.887	–1.935	–1.858	–1.719
τ^{μ} (LR)[a]	–2.615	–2.357	–2.612	–2.470
τ^{μ} (BIC)[b]	–2.205	–2.252	–2.125	–2.011
ZA^{τ} (LR)[a]	–9.799	–11.065	–9.694	–9.389
ZA^{τ} (BIC)[b]	–8.173	–12.547	–7.995	–7.557
ZA^{μ} (LR)[a]	–10.115	–11.459	–9.948	–9.952
ZA^{μ} (BIC)[b]	–8.683	–12.686	–8.441	–8.340
DF-GLS$^{\tau}$(LR)[a]	–2.653	–2.331	–2.577	–2.430
DF-GLS$^{\tau}$(BIC)[c]	–2.122	–2.545	–2.064	–2.036
DF-GLS$^{\mu}$ (LR)[a]	–2.652	–2.336	–2.570	–2.427
DF-GLS$^{\mu}$(BIC)[b]	–2.251	–2.323	–2.104	–2.090

a. Lag lengths by column (left to right) are 8, 4, 8, and 3.

b. Lag lengths by column (left to right) are 0, 1, 0, and 0.

c. Lag lengths by column (left to right) are 1, 0, 0, and 0.

NOTE: DF tests use the t-statistic for β from estimation of the equation:

$$\text{(A1)} \quad \Delta \log y_{t+1} = \alpha + \beta \log y_t + \gamma t + \sum_{i=1}^{p} \Theta_i \Delta \log y_{t+1-i} + e_{t+1},$$

τ^{τ} and τ^{μ} include and exclude the linear time trend, respectively.

ZA (usually written as Z_a) statistics are based on the same equation without lagged changes, though a choice of lag length is needed to estimate an analogue to the covariance matrix.

The τ and μ superscripts have the same significance as above.

DF-GLS$^{\mu}$ and DF-GLS$^{\tau}$, respectively, exclude and include a linear trend from the first-stage regression (as described in the text) and then estimate equation (A1) without constant, trend, or lagged changes.

LR and BIC indicate the criteria by which lag length was chosen. Critical values for τ^{τ}, ZA^{τ}, and ZA^{μ} can be found in MacKinnon (1994). For DF-GLS$^{\mu}$, the 5 percent critical value for a large number of observations is –2.89. The DF-GLS$^{\mu}$ statistic has the same distribution as τ^{μ}.

SOURCE: Author's calculations.

Perron (1995) also propose starting with a maximum value for k and decreasing the number of lags until the coefficients on the last n lagged terms are significant, but not if the total number of lags were decreased again by 1. When n=1, this procedure finds the lag length k where the t-statistic on the k+1$^{\text{th}}$ lag is not significant but the t-statistic on the k^{th} lag is significant. In his survey of the area, Stock (1994) suggests using the sequential procedure of Ng and Perron (1995), denoted *LR*, as well as the Schwartz criterion with the lag lengths constrained to between three and eight.

The unit-root tests also differ in their treatment of the "nuisance parameters" in equation (4), the constant and linear trend t. Stock (1994) evaluates the various tests and suggests using the DF-GLS test subsequently presented in detail in Elliott, Rothenberg, and Stock (1996). The DF-GLS requires two steps. First, let $z_t = (1, t)$. Then, assuming that the process for y is AR(1) with a coefficient $\alpha = 1 + c/T$, estimate $\hat{\beta}$ by regressing $[y_1, y_2(1-\alpha L), \ldots y_T(1-\alpha L)]$ onto $[z_1, z_2\,(1-\alpha L), \ldots .. z_T(1-\alpha L)]$ and then creating $y^d = y - z_t{}'\hat{\beta}$. Second, estimate equation (4) without the constant or the trend. The t-statistic on lagged y is the DF-GLS$^{\tau}$ statistic with critical values given in Elliott, Rothenberg, and Stock (1996). The DF-GLS$^{\mu}$ statistic omits t from the first stage. Its distribution is that of the more familiar Dickey–Fuller τ^{μ} statistic. The constant c is set equal to –7.0 for the no-trend case and –13.5 with trend. For comparison with Liu and Maddala, we present results of the Dickey–Fuller tests and the Phillips–Perron tests.[6] All are presented with both constant but no trend (μ) or constant and trend (τ).

III. Data and Results

We analyze data provided by Money Market Services for the deutsche mark/dollar exchange rate from January 6, 1989, through October 24, 1997.[7] The survey data represent the medians of the forecasts of the respondents for the one-week and one-month horizons (figures 1 and 2). The frequency of the data is weekly, and one month should be interpreted as corresponding to four weeks.

Table 1 indicates the unit-root test statistics for log (future DM/$) and the survey median of expectations of the future DM/$. Despite the wide variation in chosen lag lengths, in no case do we reject the null hypothesis of a unit root for either weekly or monthly forecast horizons at the familiar 5 percent level of confidence.

At the top of table 3 we also include Q-statistics similar to those presented by Liu and Maddala and which allow us to assess the extent of serial correlation present in residuals for four, eight, and 12 lags. We do this for the first differences of the individual series. Our inability to reject the null hypothesis of no serial correlation supports a conclusion of rationality.

6 The Phillips–Perron tests omit the lagged terms but still require a choice of lag length.

7 Our monthly series ends at October 3, 1997.

TABLE 2

Restricted Cointegration Test on Logarithms of Future and Expected Future Deutsche Mark/Dollar Exchange Rates: Unit-Root Tests on log (S_{t+1}) – log ($S^{e}_{t,t+1}$)

Statistics	Weekly	Monthly
τ^{τ} (LR)[a]	–7.255	–7.247
τ^{τ} (BIC)[b]	–18.979	–7.247
τ^{μ} (LR)[a]	–18.972	–7.244
τ^{μ} (BIC)[b]	–7.213	–7.244
ZA^{τ} (LR)[a]	–396.780	–103.862
ZA^{τ} (BIC)[b]	–404.361	–103.862
ZA^{μ} (LR)[a]	–398.312	–103.796
ZA^{μ} (BIC)[b]	–404.273	–103.796
DF-GLS$^{\tau}$(LR)[c]	–6.147	–5.432
DF-GLS$^{\tau}$(BIC)[b]	–18.322	–7.214
DF-GLS$^{\mu}$ (LR)[c]	–4.783	–5.315
DF-GLS$^{\mu}$ (BIC)[b]	–17.002	–7.081

a. Lag lengths by column (left to right) are 9 and 6.
b. Lag lengths by column (left to right) are 0 and 6.
c. Lag lengths by column (left to right) are 9 and 10.
NOTE: LR and BIC indicate the criteria by which lag length was chosen.
SOURCE: Author's calculations.

TABLE 3

***Q*-Statistics for Test of Serial Correlation**

Series	*Q*(4)	*Q*(8)	*Q*(12)
(1–L)* (weekly, actual)	2.1	7.3	11.2
(1–L)* (weekly, survey)	11.5	12.9	17.2
(1–L)* (monthly, actual)	2.3	7.4	11.2
(1–L)* (monthly, survey)	9.4	10.9	16.8
Forecast error, one-week forecast	6.2	10.7	16.5
Forecast error, one-month forecast[a]	7.6	13.5	20.8

a. Analysis of residual from MA(3) estimation of the weekly series of one-month-ahead forecasts.
NOTE: All $Q(k)$ statistics are distributed as x^2 (k).
SOURCE: Author's calculations.

Table 2 lists the restricted cointegration test statistics for weekly and monthly horizons. Because the test statistics are less than the critical values for the 5 percent level in all instances, we reject the null hypothesis of no cointegration. Table 3 also presents *Q*-statistics for the forecast errors. However, since the frequency of the data is weekly, a weekly series on the one-month forecast error will display serial correlation even if the forecasts are rational. In fact, this series will have a third-order moving-average error process (MA[3]).[8] In this case, we analyze the residual from estimating this series as an MA(3) process. We find no evidence of serial correlation for either of the two forecast errors.

IV. Interpretation of Results

Using more recent data and newer techniques, we confirm the findings of Liu and Maddala in favor of the rationality of Money Market Services survey forecasts. Although this would seem to support the rational expectations hypothesis, several notes of caution are in order. First, we would not necessarily confirm other implications of the rational expectations hypothesis, such as the orthogonality of forecast errors with respect to publicly available information. Second, it is not clear how median expectations are linked to the marginal prices (exchange rates) observed in the marketplace. This is one reason why economists have tended to downplay survey measures.

Nonetheless, analyses of these data bear directly on the issue of whether a risk premium exists in foreign-exchange markets and whether policy changes are anticipated correctly. Consequently, these findings are relevant to analysis of U.S. central bank intervention or other macroeconomic policies in the 1990s. Future research might fruitfully explore these possibilities.

8 We could follow Baillie and Bollerslev (1990) and impose the three moving-average parameters which would be implied by the assumption that the weekly process follows a random walk.

References

Allen, Helen, and Mark P. Taylor. "Charts, Noise and Fundamentals in the London Foreign Exchange Market," *Economic Journal,* vol. 100, no. 400 (1990), pp. 49–59.

Baillie, Richard T., and Tim Bollerslev. "A Multivariate Generalized ARCH Approach to Modeling Risk Premia in Forward Foreign Exchange Rate Markets," *Journal of International Money and Finance,* vol. 9, no. 3 (September 1990), pp. 309–24.

———, and ———. "The Forward Premium Anomaly is Not as Bad as You Think," Michigan State University, unpublished manuscript, May 1997.

Bank of Japan. *External Balance Adjustment and Monetary Policy Management under International Private Capital Flows.* Tokyo: Bank of Japan, 1989.

Beng, Gan Wee, and Wong Keng Siong. "Exchange Rate Expectations and Risk Premium in the Singapore/U.S. Dollar Exchange Rate: Evidence from Survey Data," *Applied Financial Economics,* vol. 3, no. 4 (December 1993), pp. 365–73.

Bhattacharya, Utpal, and Paul Weller. "The Advantage to Hiding One's Hand: Speculation and Central Bank Intervention in the Foreign Exchange Market," *Journal of Monetary Economics,* vol. 39, no. 2 (July 1997), pp. 251–77.

Cavaglia, Stefano, Willem F.C. Verschoor, and Christian C.P. Wolff. "Further Evidence on Exchange Rate Expectations," *Journal of International Money and Finance,* vol. 12, no. 1 (February 1993), pp. 78–98.

Dominguez, Kathryn, M. "Are Foreign Exchange Forecasts Rational? New Evidence from Survey Data," *Economic Letters,* vol. 21, no. 3 (1986), pp. 277–81.

———, and Jeffrey A. Frankel. "Does Foreign Exchange Intervention Matter? The Portfolio Effect," *American Economic Review,* vol. 83, no. 5 (December 1993), pp. 1356–69.

Elliott, Graham, Thomas J. Rothenberg, and James H. Stock. "Efficient Tests for An Autoregressive Unit Root," *Econometrica,* vol. 64, no. 4 (July 1996), pp. 813–36.

Frankel, Jeffrey A., and Kenneth A. Froot. "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations," *American Economic Review,* vol. 77, no. 1 (March 1987a), pp. 133–53.

———, and ———. "Short-term and Long-term Expectations of the Yen/Dollar Exchange Rate: Evidence from Survey Data, *Journal of the Japanese and International Economy,*" vol. 1, no. 3 (September 1987b), pp. 249–74.

Froot, Kenneth A., and Jeffrey A. Frankel. "Forward Discount Bias: Is it An Exchange Risk Premium?" *Quarterly Journal of Economics,* vol. 104, no. 1 (February 1989), pp. 139–61.

———, and ———. "Exchange Rate Forecasting Techniques, Survey Data, and Implications for the Foreign Exchange Market," International Monetary Fund, Working Paper no. 90/43, May 1990.

———, and Takatoshi Ito. "On the Consistency of Short-run and Long-run Exchange Rate Expectations," *Journal of International Money and Finance,* vol. 8, no. 4 (December 1989), pp. 487–510.

Hung, Juann H. "Intervention Strategies and Exchange Rate Volatility: A Noise Trading Perspective," *Journal of International Money and Finance,* vol. 16, no. 5 (October 1997), pp. 779–93.

Ito, Takatoshi. "Foreign Exchange Rate Expectations: Micro Survey Data," in Dilip K. Das, ed., *International Finance: Contemporary Issues.* New York: Routledge, 1993, pp. 360–74.

———. "Short-Run and Long-Run Expectations of the Yen/Dollar Exchange Rate," *Journal of the Japanese and International Economy,* vol. 8, no. 2 (June 1994), pp. 118–43.

Lewis, Karen K. "Puzzles in International Financial Markets," in Gene M. Grossman and Kenneth Rogoff, eds., *Handbook of International Economics,* vol. 3. New York: Elsevier North-Holland, 1995, pp. 1913–71.

Liu, Peter C., and G.S. Maddala. "Rationality of Survey Data and Tests for Market Efficiency in the Foreign Exchange Markets," *Journal of International Money and Finance,* vol. 11, no. 4 (August 1992), pp. 366–81.

MacDonald, Ronald, and Thomas S. Torrance. "Some Survey-Based Tests of Uncovered Interest Parity," in Ronald MacDonald and Mark P. Taylor, eds., *Exchange Rates and Open Economy Macroeconomics.* Cambridge, Mass.: Blackwell, 1989, pp. 239–48.

MacKinnon, James G. "Asymptotic Distribution Functions for Unit-Root and Cointegration Tests," *Journal of Business and Economic Statistics,* vol. 12, no. 2 (April 1994), pp. 167–76.

Ng, Serena, and Pierre Perron. "Unit Root Tests for ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag," *Journal of the American Statistical Association,* vol. 90, no. 429 (March 1995), pp. 268–81.

Schwert, William G. "Tests for Unit Roots: A Monte Carlo Investigation," *Journal of Business and Economics Statistics,* vol. 7, no. 2 (April 1989), pp. 147–59.

Stock, James H. "Unit Roots, Structural Breaks and Trends," in Robert F. Engle and Daniel L. McFadden, eds., *Handbook of Econometrics,* vol. 4. New York: Elsevier North-Holland, 1994, pp. 2740–2841.

Takagi, Shinji. "Exchange Rate Expectations: A Survey of Survey Studies," *International Monetary Fund Staff Papers,* vol. 38, no. 1 (March 1991), pp. 156–83.