SURVEY DATA ON EXCHANGE RATE EXPECTATIONS: MORE CURRENCIES, MORE HORIZONS, MORE TESTS

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Summary

This paper works with a comprehensive set of survey data, on forecasts for 24 currencies (against the dollar), thus going beyond the traditional G-7 countries to consider smaller and less developed countries. We apply the data set to four topics. (1) We find some ability in the survey data to predict future movements in the exchange rate (and in the right direction!). As in past tests, the forecasts are nevertheless biased: variability of expected depreciation is excessive, especially at the 3-month horizon. (2) We find some evidence of a time-varying risk premium, especially at the 12-month horizon. But the coefficient on the forward discount is usually significantly greater than 1/2, implying that the risk premium is less variable than is expected depreciation. (3) We examine new data on forecasts at the five-year horizon and obtain, somewhat disappointingly, only weak evidence of regressive expectations towards purchasing power parity. (4) We have no success in an attempt to use the survey data in an equation of exchange rate determination.

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1. INTRODUCTION

A new literature uses surveys of market participants to measure their expectations of exchange rates.[1] This paper extends earlier tests, based on a broader sample of survey expectations. More currencies are studied than in the early literature, many of them in newly-important emerging markets. (The countries and currencies are listed in the first appendix to the paper.) More importantly, tentative assessments are made regarding the characteristics of long horizon forecasts, whereas all prior research has focused on short to medium horizons (one week to one year). Finally, we investigate the improvement in fit for a structural exchange rate model using survey-based measures of the exchange risk premium.

The empirical results presented in this paper are based on a data set derived from Currency Forecasters' Digest (hereafter CFD). The chief advantage of exploiting this data set is that it covers over 25 exchange rates, and includes several for newly industrializing countries in Asia, and smaller developed countries in Europe and elsewhere. With a broader and more heterogeneous set of currencies, interesting patterns, not readily apparent in more narrow sets of currencies, can be identified.[2]

This paper is organized in the following fashion. In section 2, we re-evaluate the predictive ability of both the survey-based measure of expected exchange rates and the forward rate. In Section 3, the nature of the risk premium is investigated in a setting where we have direct observations on the expected depreciation, i.e., the survey measure. In section 4, the

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[1] Early contributions were Dominguez (1986), Frankel and Froot (1987), and Froot and Frankel (1989). Takagi (1991) reviewed the early literature.

[2] As was also evident in Chinn and Frankel (1994), and Frankel and Chinn (1993).

characteristics of long run forecasts are evaluated. Section 5 is composed of an investigation of the empirical usefulness of exchange rate models incorporating information about the exchange risk premium derived from survey data. Concluding remarks follow.

2. UPDATING THE STANDARD RESULTS: EXPECTATIONS AND FORWARD RATE BIAS

2.1. Overview

In these two sections we extend some of the most standard tests.[3] The exchange rate forecasts were usually compiled on the fourth Thursday of each month from February of 1988 to June of 1994.[4] The survey respondents were reported to number approximately 45, of whom two-thirds were multinational firms and the remainder forecasting firms or the economics departments of banks. We use as the measure of expectations the "consensus forecast" that CFD emphasizes. This measure is the geometric mean:[5]

$$\overline{X} \equiv \prod_{i=1}^{n} (X_i)^{w_i}$$

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[3] This is an updating results that we obtained using the Currency Forecasters' Digest data set described in Chinn and Frankel (1994) and Frankel and Chinn (1993).

[4] These data are proprietary with Currency Forecasters' Digest (acquired by the *Financial Times Group* of Alexandria, VA, toward the end of the data set) and obtained by subscription by the Institute for International Economics. The survey has apparently been conducted for some years, but the subscription of the IIE did not begin until 1988.

[5] The CFD originally described the consensus forecast as the harmonic mean. Later publications identified the consensus forecast as the geometric mean. The mode and mean responses are available, but they are very highly correlated with the geometric mean, so that the choice of measure is less likely to make much difference.

where in this case, $w_i = 1/n$,

n = number of forecasts, and

X_i represents the individual forecast responses.

The regressions are run on a pooled time series/cross section.[6] In this paper, we investigate the nature of the three- and twelve-month horizon forecasts. Since the data are sampled at intervals finer than the forecast horizon, the regression residuals will exhibit a moving average process of order k-1 (where k is the forecast horizon). Hence we use an estimate of the parameter covariance matrix that is robust with respect to heteroskedasticity and serial correlation.[7]

This new sample considerably expands the number of observations and time span investigated in previous papers. For three-month horizons, the number of observations is essentially doubled, and for 12-month, tripled. Further, the six-plus years covered by this data set encompass the 1992-93 bouts of turmoil in the EMS (European Monetary System), the depreciation of the dollar against the yen in the early 1990s (see Figures 1 and 2), and the acceleration in capital flows to emerging markets that preceding the Mexican peso crisis of December 1994 and the East Asia crisis of 1997-98 .

[6] We also ran regressions on individual time series (reported in Appendices 1-3). The results are consistent in a qualitative sense with those reported in the text, although there is much variation in the estimated slope coefficients, as one would expect from the relatively small number of observations in each time series.

[7] The robust standard errors take into account the moving-average serial correlation induced by overlapping horizons, but do not adjust for cross-currency correlation. In tests on the entire sample, regression residuals do not exhibit substantial cross-currency correlation. (The correlation for regression (1) residuals is about 0.13.)

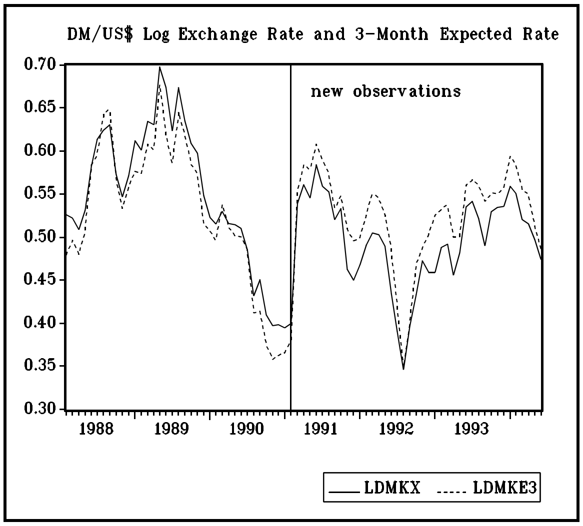

Figure 1

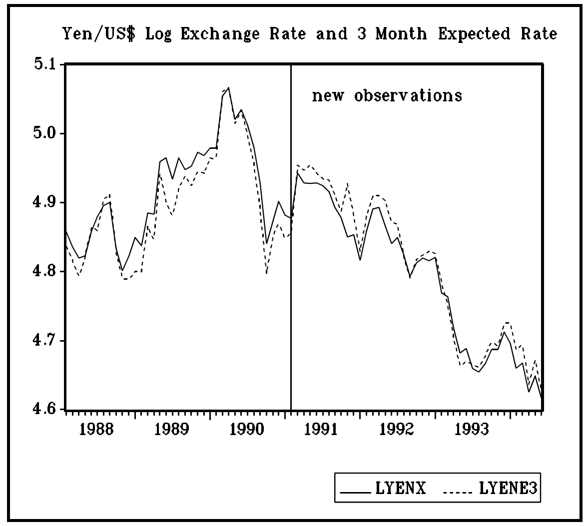

Figure 2

2.2. Conditional Bias in Expectations

We now explicitly evaluate of the forecasting characteristics of our expectations measures. A common procedure is to regress the ex post depreciation on the survey measure of expected depreciation. In that case, the unbiasedness proposition is represented by the null hypothesis that the coefficient on expected depreciation equals unity.

$$\Delta s_{t+k} = \alpha_1 + \beta_1 \Delta \hat{s}^{e}_{t,t+k} + u_{1,t} \tag{1}$$

where Δs_{t+k} is the ex post depreciation between periods t and t+k, and $\Delta\hat{s}_{t,t+k}^{e}$ is the period t survey-based expected depreciation from time period t to t+k (all variables are annualized). The unbiasedness hypothesis is represented by the null hypothesis that $\beta_1=1$, and the alternative, by $\beta_1 < 1$. If the alternative is accepted, then investors could make better guesses by betting against the consensus forecast. These regressions were run on both the entire sample, and a sample excluding the three countries that historically had had high inflation rates, Argentina, Brazil and

Mexico. Estimates for this equation are presented in Table 1.

Results for six regressions are reported. Two constrain the intercept for all currencies to be the same. Since such constraints are usually rejected statistically, the constrained regression results are provided only for purposes of comparison. Focusing attention on the full sample-unconstrained regressions, one finds that the estimate of β_1 is usually positive and large in economic terms. One can reject the null hypothesis that the coefficient is zero at the three-month horizon, and can reject it at the 10% level at the 12-month horizon. In the three-month case the null of a unit coefficient is strongly rejected, indicating bias in expectations. On the other hand, when the intercepts are unconstrained at the 12-month horizon, unbiasedness is not rejected.

It appears here (as in our previous papers) that the results, particularly the finding of some predictive ability, are being driven by the inclusion of currencies for high inflation countries. When the sample excludes the currencies of Argentina, Brazil and Mexico, coefficient estimates become either negative (at the 3-month horizon) or insignificantly different from zero. In these cases, the null of unbiasedness is resoundingly rejected.

This finding of conditional bias over the 1988-94 period corroborates results in Frankel and Froot (1987), which pertained to a turbulent period in the 1980s encompassing a large rise and fall in the value of the dollar. Here, the results obtain over a period of relative dollar stability against the European currencies, combined with uncertainty within the EMS.

The finding of parameter estimates substantially below unity may be due to the presence of measurement error in the survey forecasts. Even if such measurement error is randomly distributed, the coefficient estimate in equation (1) is biased downward.[8]

2.3. Forward Rate Bias

Many studies have concluded that the forward discount is a biased predictor of the future spot rate. Below, we replicate the conventional test for forward rate bias using our data sample:

$$\Delta s_{t+k} = \alpha_2 + \beta_2 fd_{t,t+k} + u_{2,t} \tag{2}$$

where $fd_{t,t+k}$ is the forward discount at time t for a trade k periods hence.

The results of estimating this equation are reported in Table 2 (only 17 of the 24 currencies are covered since not all currencies have forward markets). One can view this regression as an explicit test of the unbiasedness properties of the forward discount. In contrast to some estimates in this literature, as well as in our own earlier findings, we find that the point estimates for the forward discount are positive. Moreover, the null of unbiasedness cannot be rejected in the cross section regressions.

These surprising findings are apparently being driven by the behavior of those European currencies that were forced to devalue in the crises of 1992 and 1993 in the ERM (Exchange Rate Mechanism). A regression using only EMS currencies yields point estimates (GMM standard errors) of 1.214 (0.553) and 1.606 (0.461) for the three-month and 12-month

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[8] Attempts to address this problem by application of two stage least squares, using lagged values of expected depreciation as instrumental variables, usually reinforce the results obtained above. When two-stage least squares is applied to a regression of ex post depreciation on expected depreciation, the coefficients become even more negative. Because the standard errors become larger, however, rejection of the null hypothesis of unity becomes more difficult.

horizon regressions.[9] This difference in slope parameters implies that the non-EMS currencies have lower coefficients. In fact, a cross-section regression for non-European currencies yields estimates of -1.483 (0.594) and -1.524 (0.763).

Moreover, inspection of the time series regressions bears out the assertion (see Appendix 2). The currencies with coefficients of 1.0 or greater are the Swedish krone, Irish punt, Italian lira, Norwegian krona, and Spanish peseta. Hence the forward discount appears during this period in some degree to have correctly anticipated the subsequent movements in the European currencies. This could be an instance of the much-discussed "peso problem."[10] It could also be evidence that, after twenty years of observed bias in the forward discount, market participants began to adjust their behavior. The "convergence play" of the early 1990s, when speculators went long in those European currencies that were paying high interest rates, would be an example. A further-updated test confirms that the coefficients on two major European currencies that showed a large bias during the period 1976I-1989IV, showed a smaller bias during the period 1990I-1998I.[11]

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[9] Flood and Rose (1994) also find that the forward discount puzzle does not hold for EMS currencies. In other words, the forward rate correctly predicts the direction of change in the spot rate in these cases.

[10] The peso problem is usually invoked (often ex post) to explain forecasts of changes in the exchange rate that did not materialize, or materialized to a smaller extent than forecast. The crude explanation is that "the feared devaluation did not happen to occur during the sample period." More correctly, the non-normality of the distribution produces biased standard errors by which to evaluate the significance. If peso problems abound, however, *some* studies should catch a sample period when the devaluation *did* happen to occur, and thus when the change in the exchange rate materialized to a greater extent than forecast. This appears to describe the European currencies in our sample.

[11] Not reported in the tables. The estimated coefficient changes from -2.7 to -0.5 for Germany, and from -2.2 to 1.2 for the United Kingdom. For France and Italy, there was less bias in the 1970s and 1980s to begin with. The tests use the 3-month offshore interest

Most other currencies in our sample show the familiar pattern of small or negative coefficients. The yen, Australian dollar (12 month), Singapore dollar, South African Rand and Canadian Dollar all show statistically significant deviations from unbiasedness. The reason that the deviations for other currencies are not significant is presumably the large standard errors attendant with these pure time series regressions.

3. THE FORWARD DISCOUNT AND THE EXCHANGE RISK PREMIUM

The forward rate bias found for certain currencies could be due either to variation in the risk premium or to within-sample bias in expectations. Recall equation (2):

$$\Delta s_{t+k} = \alpha_2 + \beta_2 fd_{t,t+k} + u_{2,t}$$

The null hypothesis of unbiasedness is represented as $ß_2=1$. (A constant is allowed to account either for a constant risk premium, or for the convexity term arising from Jensen's Inequality.) A common finding in the literature is rejection of the null, where $ß_2$ is usually estimated to be closer to zero than to unity. This finding is taken to be evidence that most of the variation in the forward discount constitutes a time-varying risk premium (defined by $rp_{t,t+k} \equiv fd_{t,t+k} - \Delta s^e_{t,t+k}$).

As discussed above, the forward rate appears to be a biased predictor of the spot rate for particular currencies, and not for others. While the null of unbiasedness cannot be rejected for all currencies in general, in part because the standard errors are quite large, we have seen that

differential and heteroscedasticity-consistent standard errors. But the data are quarterly, so the standard errors are fairly large (2.1 and 1.7 for Germany; 0.9 and 1.2 for the UK).

unbiasedness is clearly rejected for non-European currencies. The question is what is the source of the bias in the forward discount.

To assess whether the bias is due to expectational errors or a time-varying risk premium, one can regress the expected depreciation, $\Delta\hat{s}^e$ as estimated by the CFD survey, on the forward discount, as suggested by Froot and Frankel (1989). That is:

$$\Delta\hat{s}^{e}_{t,t+k} = \alpha_3 + \beta_3 fd_{t,t+k} + u_{3,t} \tag{3}$$

The results are reported in Table 3. The null hypothesis that the slope coefficient is zero is strongly rejected.[12] Thus, at least some of the variation in the forward discount must be due to expected depreciation. In other words, one can reject the hypothesis that all of the variation in the forward discount is due to a time-varying risk premium.[13] This finding also supports the validity of the survey data as a measure of expected depreciation, as opposed to the null hypothesis that the survey responses are random noise.

The next question is whether any of the variation in the forward discount can be attributed to a risk premium or, in other words, whether we can reject the hypothesis of a unit coefficient. Here we get somewhat different answers depending on whether we look at the three-month results or the 12-month results. Overall, there is more evidence to support the existence of a risk premium in this cross-section of 17 currencies than there was in the earlier studies of five

[12] Results of these regressions report GMM standard errors since there is some evidence of serial correlation. Although the correlation is not due to overlapping observations, empirically, the assumption of MA(2) errors in calculating robust standard errors appears adequate. The assumption of higher-order MAs yielded similar estimates of the standard errors.

[13] The same is true for most of the currencies considered individually (Appendix Table A3), despite the much-reduced degrees of freedom.

major currencies. The coefficient estimate of .736 is significantly different from 1, for example, in the case where the 12-month horizon is used and the intercept terms are constrained to be equal across currencies.[14]

At the three month horizon one can reject the null hypothesis that $\beta_3 = 1$. The rejection however is in the direction $\beta_3 > 1$, which is inconsistent with the usual view that some positive share of the variation in the forward discount constitutes variation in the risk premium.[15]

The regression is also capable of shedding light on a claim set forth by Fama (1984) and Hodrick and Srivastava (1986) (FHS) that expected depreciation is less variable than the exchange risk premium. The FHS claim is:

$$var(\Delta s^{e}_{t,t+k}) < var(rp_{t,t+k}) \tag{4}$$

To see the relevance of the regression results for this claim, note that (4) can be re-written (dropping subscripts) as:

$$var(\Delta s^{e}) < var(fd) + var(\Delta s^{e}) - 2 \times cov(fd, \Delta s^{e})$$

Rearranging:

[14] This can be considered a more powerful test of the no-risk-premium hypothesis than the unconstrained case, because under that null hypothesis all intercept terms are zero. The rejection is even stronger when the sample is restricted to non-European currencies.

[15] When the European currencies are excluded, the coefficient is very close to 1, suggesting no time-varying risk premium, in line with the results from the narrower 5-currency sample of Froot and Frankel (1989).

$$\frac{1}{2} \geq \frac{cov(fd,\Delta s^{e})}{var(fd)} \tag{5}$$

As Froot and Frankel (1989) observe, the probability limit of the ß coefficient in (3) is:

$$plim\,\hat{\beta}_3 = \frac{cov(u_3,fd) + cov(\Delta\hat{s}^{e},fd)}{var(fd)} \tag{6}$$

Assuming the measurement error is uncorrelated with the forward discount, then the probability limit of the regression estimate is the same as the expression in the RHS of (5). Hence, if one can reject the null hypothesis that $\beta_3 \leq 0.5$, then one is rejecting the FHS hypothesis that the variation in the expectation of depreciation is less than the variation in the risk premium.

In almost all cases, one can reject the hypothesis $\beta_3 \leq 0.5$, indicating there is less evidence of a time-varying risk premium than in the previous results reported in Frankel and Chinn (1993). However, at the 12-month horizon for the non-European currencies, the null hypothesis $\beta_3 \leq 0.5$ is not rejected, implying some role for a risk premium.

4. LONG-HORIZON FORECASTS

4.1. Extrapolative and Adaptive Expectations

In this section we evaluate the characteristics of multi-year forecasts provided by CFD. Previous studies used, at the longest horizon, 12 month ahead forecasts. In contrast we examine the behavior of five-year ahead forecasts. We evaluate the evidence in favor of extrapolative, adaptive and regressive expectations in a variety of cross-section regressions.

CFD reports the exchange rate expected four to five years into the future. For instance, the March 1993 issue reports forecasts for the end of 1993, 1994, 1995, 1996, and 1997; the

June, September and December 1993 issues, also report forecasts for the end of these same years. Hence the forecasts are not precisely 5 years-ahead, except in the case of the March 1993 issue. To be safe, we interpret only data for the first observation of the year as a 5-year-ahead forecast. This produces 6 observations per currency, or a total of approximately 150 observations in the cross-section sample.

In Table 4.1, results are reported for regressions of the form:

$$\Delta\hat{s}^{e}_{t,t+60} = \alpha_4 + \beta_4 \Delta s_{t,t-k} + u_{4,t} \tag{7}$$

where k is either 1 or 3. If there are bandwagon effects, then the coefficient estimate should be positive. On the other hand, if expectations are stabilizing, the coefficient estimate should be negative. With static expectations (which would be consistent with the stylized fact of random walk exchange rates), the coefficient should be zero.

The point estimate is significant and positive, which is indicative of bandwagon effects. At the end of a month when the currency has depreciated at a rate of 1 per cent, the expectation is for a further depreciation over the next five years at a rate of about .05%. (All variables are at annualized rates.[16]) However the standard errors are sufficiently large to make these estimates statistically insignificant, except when the intercept is unconstrained. Moreover, this point estimate seems to be driven by the inclusion of the three currencies that during the sample period were coming off of very high inflation rates -- Argentina, Brazil and Mexico. Omitting these observations leads to a negative estimate. Slightly more evidence for bandwagon

[16] In absolute terms, the estimated effect of a 1 % depreciation is an expected cumulative 3 % further depreciation [=.05×5×12].

effects is found when using 3-month lagged depreciation. However the significance of these findings is again not robust to the omission of the high-inflation currencies.

Another possible expectations formation scheme is adaptive. We model adaptive expectations according to the following equation:

$$\Delta \hat{s}^{e}_{t,t+60} = \alpha_5 + \beta_5 (\hat{s}^{e}_{t-12,t} - s_t) + u_{5,t} \tag{8}$$

The results of running this regression, using 12 month expectations, are reported in Table 4.2. Surprisingly, regardless of the specification, the estimated coefficient is negative. A negative coefficient implies that agents heavily discount year-old forecasts of the current exchange rate in forming their expectations regarding the exchange rate 5 years hence, and rely instead on the most recent spot rate. Once again, much of the result appears driven by the inclusion of the high inflation currencies. Nonetheless, the estimate is statistically negative even in the low-inflation sample (although it too is economically small).

These two sets of results suggest that even for horizons far in the future, expectations may not be stabilizing. Rather it appears that agents place a great deal of weight on the current spot rate in formulating their expectations regarding even the very remote future. One cannot make strong conclusions regarding the stabilizing attributes of expectations, however, without regard to some sort of fundamentals.

4.2. Regressive Expectations

In certain cases, regressive expectations may be the same as rational expectations, as in Dornbusch (1976). A reasonable interpretation of regressive expectations is:

$$\Delta s^{e}_{t,t+k} \;=\; -\upsilon(s_t - \overline{s}_t) \tag{9}$$

where the over-bar indicates the long run equilibrium value. One can investigate whether the long-run value of the nominal exchange rate is constant, or whether it varies with price levels according to PPP. This suggests two regressions:

$$\Delta\hat{s}^{e}_{t,t+60} \;=\; \alpha_6 \;+\; \beta_6 s_t \;+\; u_{6,t} \tag{10}$$

; and

$$\begin{aligned}\Delta\hat{s}^{e}_{t,t+60} \;&=\; \alpha_7 \;+\; \beta_7(s_t \;-\; \overline{s}_t) \;+\; u_{7,t} \\ where \quad \overline{s}_t \;&\equiv\; \log(CPI^{\ell}_t / CPI^{US}_t)\end{aligned} \tag{11}$$

for the case of a constant long run real exchange rate. (The currency dummies pick up the base year effects.) In both cases, the slope coefficient should be negative under the hypothesis that the exchange rate is expected to regress back toward equilibrium. Note that if inflation rates are always the same in the two countries, then $ß_6 = ß_7$.

The results of these regressions are reported in Table 4.3. The estimates of the slope coefficient in equation 10 are positive, but the standard errors are sufficiently large that the 90% confidence bounds easily encompass negative values.

The assumption of a constant long run nominal exchange rate is implausible for currencies experiencing very different price trends. The regression for equation 11 over the entire sample, allowing for a long run exchange rate moving with PPP, yields a negative, but statistically insignificant estimate of -0.029.

Some of the currencies included in this sample experienced fairly rapid inflation. As Frankel (1979) points out, in the presence of secular inflation, equation 9 is no longer appropriate, and should be rewritten as:

$$\Delta s^{e}_{t,t+k} \;=\; -\upsilon(s_t - \bar{s}_t) \;+\; \pi^{\ell}_t \;-\; \pi^{US}_t \tag{12}$$

where the inflation rates are the expected rates over the next five years. In the absence of expectations data regarding inflation rates, we proxy them by using 3 month and 12 month lagged ex post inflation. Re-estimating equation 11 over the same sample, allowing for inflation differentials:

$$\begin{aligned} \Delta\hat{s}^{e}_{t,t+60} \;&=\; \alpha_7 \;+\; \beta_7(s_t \;-\; \bar{s}_t) \;+\; \gamma_7(\pi^{\ell}_t \;-\; \pi^{US}_t) \;+\; u_{7,t} \\ \textit{where} \;\; \bar{s}_t \;&\equiv\; \log(CPI^{\ell}_t/CPI^{US}_t) \end{aligned} \tag{13}$$

produces a slightly more negative point estimate, but results that are still not statistically significant. (Results using 3 month inflation rates are reported in Table 4.3, but the results using 12 month rates do not differ substantially.)

This particular specification relies upon PPP holding in the long run. If the PPP condition fails to hold even in the long run, perhaps due to the presence of nontradables and differential productivity growth rates, then one might expect a better fit for a sample that excludes currencies associated with Asian countries that are thought to have experienced high productivity growth. We report results for a sample including only the European currencies, in the right-most column of Table 4.3. In this sample we do obtain an economically significant negative coefficient estimate, as hypothesized (although statistically significant only at the 20%

significance level). The point estimate, -0.059, implies that if the currency today is overvalued by 10 per cent according to the PPP criterion, market participants expect it to depreciate at an annual rate of 0.59 per cent over the next five years, or 3.0 per cent cumulatively, a plausible estimate.

5. SURVEY DATA AND STRUCTURAL MODELS OF EXCHANGE RATES

One of the most resilient stylized facts of international finance is the failure of structural models of exchange rate determination to fit the data well, particularly out of sample (e.g., Meese and Rogoff, 1983). Instead of looking to new econometric techniques to gain insight into exchange rate dynamics, we investigate here whether including a time-varying risk premium, as measured by the survey data, improves in-sample fit.

Assume in a conventional monetary model (perhaps with sticky prices) covered interest parity holds as usual,

$$i^{\ell}_{t,t+k} - i^{US}_{t,t+k} = f_{t,t+k} - s^{e}_{t,t+k}$$

but the uncovered interest parity condition is modified by the addition of a risk premium:

$$i^{\ell}_{t,t+k} - i^{US}_{t,t+k} - \Delta s^{e}_{t,t+k} = rp_t$$

This leads to an expression for the exchange rate of the form:

$$\begin{aligned} s_t &= \gamma_0 + \gamma_1 \hat{m}_t + \gamma_2 \hat{y}_t + \gamma_3 \hat{i}_t + \gamma_4 \hat{\pi}_t + \gamma_5 rp_t \\ &\quad \text{where} \quad rp_t \equiv f_{t,t+k} - \hat{s}^{e}_{t,t+k} \end{aligned} \tag{14}$$

where the carats "^" indicate relative differences (local minus US). The coefficients on money and inflation should be positive, and those on income and interest rates negative. The coefficient on the risk premium should be positive, and equal in magnitude to the negative of the interest rate coefficient.[17]

Monthly data on money (M1), income (industrial production), interest rates (usually interbank), and 3-month CPI inflation rates were drawn from the IMF's International Financial Statistics database. All these macroeconomic series are seasonally adjusted. The risk premium was calculated using the forward rate and survey data on expectations at the 3-month and 12-month horizon, used in the previous analyses. Of course, we do not observe the true risk premium, merely the risk premium implied by the survey data.

The regressions were run on the pooled time-series/cross-section for the period 1988.04 to 1994.06. Some currencies were omitted, either because they lacked forward markets or because certain macro data were unavailable. In total, 13 of the possible 17 currencies were included in the regressions.

The results of estimating equation (14) in first differences are reported in Table 5.[18] In the first column are the estimates for the standard specification, without country dummies. The proportion of variance explained is very low, and only income shows up significantly, and in the correct direction. In the second column, each currency is allowed a separate intercept, with no appreciable change in the results. In the next column, the 3-month risk premium is included. It

[17] This equation was derived and tested, for example, as a synthesis of the monetary and portfolio-balance models (equation 25), in Frankel (1983).

[18] The regressions are implemented in first differences because the levels regressions exhibited extreme serial correlation.

enters in negatively and with statistical significance. The 12-month risk premium is included in the last column, and enters without statistical significance. The results using long term interest rates instead of short term are not appreciably different.

The sign on the risk premium (or, equivalently, on expected depreciation) is the opposite from that hypothesized. It is possible that the problem is simultaneity bias; after all, in Section 4 of the paper we estimated equations in which expectations depended on the spot rate rather than the other way around. Further research could use news reports of central bank intervention as an instrumental variable. For the time being, we must conclude that our approach to entering expectations as an explanatory factor is no more encouraging than other typical equations of exchange rate determination.

6. CONCLUSIONS

We have applied a comprehensive data set, consisting of forecasts for 24 currencies against the dollar, to four topics: (1) the ability to predict the spot rate, (2) evidence of a time-varying risk premium, (3) the formation of expectations, and (4) exchange rate determination.

On the first topic, we have found some predictive power in the survey data (and in the right direction!).[19] As in past tests, the forecasts are nevertheless biased: variability of expected depreciation is excessive (in the sense of Bilson, 1981), especially at the 3-month horizon.

On the second topic, we have found some evidence of time-varying risk premium,[20]

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[19] Results in Chinn and Frankel (1994, p.762) show some predictive power, but usually in the wrong direction (especially if Argentina, Brazil, and Mexico are excluded).

[20] As in Frankel and Chinn (1993).

especially at the 12-month horizon. But the coefficient on the forward discount is usually significantly greater than 1/2, implying that the $var(rp) < var(\Delta s^e)$.[21]

On the third topic, we examined new data on forecasts at the five-year horizon, and found weak evidence of regressive expectations. The estimate implies that if a European currency today is overvalued by 10 per cent according to the PPP criterion, market participants expect it to depreciate 3 per cent cumulatively over the next five years. The result is consistent with past findings that, as the horizon lengthens from one week to three months to one year, expectations become increasingly regressive.[22]

On the fourth topic, we have so far had no more success than anyone else. Simultaneity is a likely culprit.

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[21] As in Froot and Frankel (1989).

[22] E.g., Frankel and Froot (1987).

DATA APPENDIX

Currency Forecasters' Digest is published monthly. The publication indicates that the forecasts apply to a specific date, usually either the third or fourth Thursday in the month. The forecasts include 1, 3, 6 and 12 month horizon forecasts, with the following measures: Geometric mean, arithmetic mean and modal mean. Contemporaneously dated spot rate data are also provided. All rates are converted to domestic currency units per US dollar.

The following currencies are surveyed:

Mnemonic	Currency	FR	A	T/I	E
DM	West German DM	F			E
FFR	French Franc	F			E
DKR	Danish Krone	F			E
UK	UK Pound Sterling	F			E
NTH	Netherlands Guilder	F			E
SFR	Swiss Franc	F			
SKR	Swedish Krone	F			
IRE	Irish Punt	F			E
BFR	Belgian Franc	F			E
LIR	Italian Lire	F			E
NKR	Norwegian Krone	F			
SP	Spanish Peseta	F			
YEN	Japanese Yen	F			
TAI	Taiwanese Dollar				
AUS	Australian Dollar	F			
SNG	Singapore Dollar	F	A		
PHL	Philippine Peso		A		
KOR	Korean Won				
SAR	South African Rand	F	A		
CAN	Canadian Dollar	F			
ARG	Argentine Austral			I	
MEX	Mexican Peso			I	
CHL	Chilean Peso			T	
BRZ	Brazilian Cruzeiro/ado			I	
BOL	Venezuelan Bolivar		A		

Key: F: Forward rate available. A: Alternating monthly. T: series terminates before June 1994. I: Many missing values due to currency change. E: EMS currency categorization, for this paper.

Forward rates are the arithmetic average of bid and ask rates at London close, as reported by DRIFACS.

To minimize the number of missing observations, a recursive Chow-Lin (1976) procedure for interpolation was used for the expectations series. The missing observations are November 1989, February 1990 and April 1990. The related series used in the interpolation procedure is the contemporaneous (log) spot rate.

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TABLE 1

Regression of actual depreciation on expected depreciation

February 1988 to June 1994

$$\Delta s_{t+k} = \alpha_1 + \beta_1 \Delta \hat{s}^{e}_{t,t+k} + u_{1,t} \qquad (1)$$

Term (k)	3 month (intercept constrained)	3 month	3 month (No Arg., Brz.,Mex.)	12 month (intercept constrained)	12 month	12 month (No Arg., Brz.,Mex)
$ß_1$	0.740	0.461	-0.163	1.411	0.828	0.003
SE	(0.017)	(0.029)	(0.058)	(0.025)	(0.045)	(0.057)
GMM SE	(0.077)	(0.123)	(0.096)	(0.303)	(0.453)	(0.115)
t: $ß_1$=1	3.377***	4.382***	12.115***	1.356	0.380	8.670***
t: $ß_1$=0	9.610***	3.748***	1.698*	4.657***	1.828*	0.026
df	1639	1616	1409	1458	1435	1287
$\bar{R}^2$	0.54	0.58	0.02	0.68	0.74	0.14
DW	0.514	0.456	0.566	0.143	0.105	0.248

Notes: "Intercept constrained" indicates that all the exchange rates are constrained to have the same intercept term. OLS $\hat{ß}$ is the point estimate from the OLS regression. SE is the OLS asymptotic standard error. GMM is a heteroskedasticity and serial correlation consistent-Generalized Method of Moments standard error. GMM SE is from regressions with de-meaned data. *(**)[***] indicates significance at 10% (5%) [1%] level.

TABLE 2
Bias in the Forward Discount
February 1988 to June 1994

$$\Delta s_{t+k} = \alpha_2 + \beta_2 fd_{t,t+k} + u_{2,t} \tag{2}$$

Term (k)	3 month constr.	3 month	3 month EMS	12 month constr.	12 month	12 month EMS
OLS $\hat{\beta}_3$	0.883	0.805	1.214	1.142	1.113	1.606
OLS SE	(0.169)	(0.209)	(0.311)	(0.101)	(0.140)	(0.190)
GMM SE	(0.327)	(0.382)	(0.346)	(0.346)	(0.410)	(0.461)
t: β_2=0	2.700***	2.107***	2.195**	3.301***	2.715***	3.484***
t: β_2=0.5	1.171	0.798	1.291	1.855**	1.226	2.399**
t: β_2=1	0.358	0.301	0.387	0.410	0.268	1.315
d.f.	1246	1230	680	1127	1111	609
$\bar{R}^2$	.02	.01	.01	.10	.10	.12
DW	0.548	0.559	0.550	0.259	0.224	0.230

Notes: OLS ß is the point estimate from the OLS regression. SE is the standard error. OLS Het. SE is the White heteroskedasticity consistent SE. GMM SE is a heteroskedasticity-consistent Generalized Method of Moments standard error. GMM SE is from regressions with de-meaned data. t is the absolute value of the t-statistic using the OLS point estimate and GMM standard error. *(**)[***] indicates significance at 10% (5%) [1%] level.

TABLE 3
Test for Time Varying Risk Premium
February 1988 to June 1994

$$\Delta\hat{s}^{e}_{t,t+k} = \alpha_3 + \beta_3 fd_{t,t+k} + u_{3,t} \tag{3}$$

Term (k)	3 month constr.	3 month	3 month constr., Non-Europe	12 month constr.	12 month	12 month constr., Non-Eur.
OLS $\hat{\beta}_3$	1.403	1.820	0.956	0.736	1.126	0.445
SE	(0.069)	(0.081)	(0.095)	(0.042)	(0.050)	(0.062)
GMM SE	(0.111)	(0.165)	(0.171)	(0.068)	(0.084)	(0.105)
t: β_3=0	12.640***	11.030***	5.591***	10.824***	13.405***	4.238***
t: β_3=0.5	8.135***	9.212***	2.667***	3.471***	7.452***	0.524
t: β_3=1	3.631***	4.976***	0.257	3.882***	1.500	5.286***
d.f.	1230	1212	315	1231	1213	316
$\bar{R}^2$	.25	.30	.24	.20	.38	.14
DW	0.574	0.707	0.852	0.516	0.715	0.375

Notes: OLS ß is the point estimate from the OLS regression. SE is the standard error. OLS Het. SE is a White heteroskedasticity-consistent standard error. GMM is a heteroskedasticity consistent-Generalized Method of Moments standard error, assuming MA processes of order two. Assuming higher order lags implies only slightly different results. t is the absolute value of the t-statistic using the OLS point estimate and either the White heteroskedasticity-consistent or the GMM standard error.
* (**)[***] indicates significance at 10% (5%) [1%] level.

TABLE 4.1
Extrapolative Expectations in Long Horizon Forecasts
February 1988 to June 1994

$$\Delta\hat{s}^{e}_{t,t+60} = \alpha_4 + \beta_4 \Delta s_{t,t-k} + u_{4,t} \tag{7}$$

Term (k)	1 month lagged depreciation			3 month lagged depreciation		
	constrained		No ABM	constrained		No ABM
OLS $\hat{\beta}_4$	0.044	0.040**	-0.006			
SE	(0.040)	(0.016)	(0.009)			
OLS $\hat{\beta}_4$				0.206***	0.176***	0.026
SE				(0.045)	(0.040)	(0.028)
d.f.	129	106	98	129	106	121
$\bar{R}^2$	.02	.86	.52	.06	.56	.52

Notes: OLS ß is the point estimate from the OLS regression. SE is the OLS standard error, assuming T/3 observations. t is the absolute value of the t-statistic using the OLS point estimate and adjusted standard error.
*(**)[***] indicates significance at 10% (5%) [1%] level.

TABLE 4.2
Adaptive Expectations in Long Horizon Forecasts

February 1988 to June 1994

$$\Delta\hat{s}^{e}_{t,t+60} = \alpha_5 + \beta_5(\hat{s}^{e}_{t-12,t} - s_t) + u_{5,t} \tag{8}$$

Term (k)	12 month forecast error constrained		No ABM
OLS $\hat{\beta}_5$	-0.100	-0.091	-0.090**
SE	(0.102)	(0.086)	(0.040)
d.f.	132	120	103
$\bar{R}^2$	.27	.45	.48

Notes: OLS ß is the point estimate from the OLS regression. SE is the OLS standard error, assuming T/3 observations. t is the absolute value of the t-statistic using the OLS point estimate and adjusted standard error. *(**)[***] indicates significance at 10% (5%) [1%] level for the null hypothesis of zero coefficient.

TABLE 4.3
Regressive Expectations in Long Horizon Forecasts
February 1988 to June 1994

Term (k)	Constant LR Exchange Rate	PPP Varying Long Run Exchange Rate		
	No ABM	No ABM	No ABM	Europe
OLS $\hat{\beta}_6$	0.0098			
SE	(0.045)			
OLS $\hat{\beta}_7$		-0.0126	-0.0285	-0.0590
SE		(0.037)	(0.044)	(0.046)
OLS $\hat{\gamma}_7$			0.124***	-0.006
			(0.054)	(0.063)
d.f.	120	110	100	64
SER	2.762	3.111	2.758	2.078
$\bar{R}^2$	.52	-.01	0.04	-.01

Notes: Regression estimates from equations 10, 11 and 13. In this table, depreciation rates are expressed in decimal form (i.e., 10% is .10) so that coefficients are interpretable as elasticities. SE is the OLS standard error. t is the absolute value of the t-statistic using the OLS point estimate and adjusted standard error. *(**)[***] indicates significance at 10% (5%) [1%] level for the null hypothesis of zero coefficient. In regressions involving PPP, the following currencies were omitted: IRE, SFR, AUS and BOL.

TABLE 5
Structural Models of the Exchange Rate
April 1988 to June 1994

Coeff.	[1]	[2]	[3]	[4]
m	-0.0019 (0.020)	-0.0011 (0.025)	-0.0005 (0.025)	0.0006 (0.025)
y	-0.0662* (0.039)	-0.0653* (0.039)	-0.0818** (0.042)	-0.0789* (0.042)
i (x100)	0.0088 (0.034)	0.0091 (0.034)	0.0244 (0.035)	0.0127 (0.035)
π (x100)	0.0058 (0.049)	0.0595 (0.050)	0.0709 (0.052)	0.0763 (0.052)
rp3 (x100)			-0.0717*** (0.017)	
rp12 (x100)				-0.0528 (0.038)
currency dummies	no	yes	yes	yes
d.f.	903	881	823	825
$\bar{R}^2$	.00	-.01	.01	-.01
DW	1.789	1.799	1.879	1.829

Notes: Estimates from first difference regressions. OLS standard errors in parentheses. Interest and inflation rate, risk premia coefficients multiplied by 100. Currencies covered: DMK, FFR, DKR, UKM, NTH, SKR, BFR, LIR, NKR, SPN, YEN, SAF, CAN.

TABLE A1
Tests of Bias in Survey Expectations
February 1988 - June 1994

Exch Rate	Term (k)	Const.	β_1	$\bar{R}^2$	SER	DW	d.f.
DMK	3	-0.989 [5.261]	-0.307 [0.384]	0.01	24.679	0.553	75
DMK	12	-1.231 [4.400]	-0.043 [0.461]	-.01	10.318	0.238	67
FFR	3	-0.459 [5.703]	-0.222 [0.415]	-.00	24.202	0.508	75
FFR	12	-0.063 [5.416]	-0.184 [0.554]	-.01	11.002	0.213	67
DKR	3	-0.387 [5.751]	-0.226 [0.438]	-.00	25.054	0.513	75
DKR	12	-0.086 [5.562]	-0.156 [0.545]	-.01	11.636	0.211	67
UK	3	3.478 [5.367]	-0.309 [0.348]	0.00	27.455	0.613	75
UK	12	3.885 [2.341]	-0.198 [0.556]	-.01	11.657	0.257	67
NTH	3	-0.872 [5.391]	-0.287 [0.401]	-.01	24.842	0.548	75
NTH	12	-1.297 [4.390]	-0.041 [0.449]	-.01	10.355	0.238	67
SFR	3	-0.104 [5.723]	-0.497 [0.437]	0.03	26.881	0.542	75
SFR	12	1.074 [5.749]	-0.429 [0.540]	0.02	12.091	0.226	67

Exch Rate	Term (k)	Const.	β_1	$\bar{R}^2$	SER	DW	d.f.
SKR	3	3.448 [5.760]	0.0346 [0.321]	-.01	29.463	0.465	75
SKR	12	0.795 [2.554]	0.544 [0.506]	0.03	14.192	0.173	67
IRE	3	-0.258 [5.552]	0.105 [0.385]	-.03	21.870	0.460	75
IRE	12	-0.583 [6.444]	0.132 [0.411]	-.01	12.286	0.182	69
BFR	3	-1.262 [5.657]	-0.125 [0.402]	-.01	25.371	0.514	75
BFR	12	-0.532 [5.248]	-0.159 [0.526]	-.01	11.310	0.219	69
LIR	3	2.503 [5.828]	0.125 [0.417]	-.01	28.335	0.550	75
LIR	12	-0.471 [4.475]	0.529 [0.686]	0.02	13.254	0.185	67
NKR	3	2.162 [5.189]	-0.055 [0.253]	-.01	24.330	0.581	66
NKR	12	0.930 [3.013]	0.159 [0.354]	-.01	10.243	0.250	62
SP	3	1.158 [5.450]	0.107 [0.384]	-.01	27.481	0.568	75
SP	12	-0.721 [3.792]	0.604 [0.461]	.04	13.901	0.197	67
YEN	3	-5.066 [3.859]	-0.726 [0.326]	0.10	20.572	0.682	75
YEN	12	0.367 [3.090]	-1.352 [0.269]	0.39	7.475	0.610	67

Exch Rate	Term (k)	Const.	β_1	$\bar{R}^2$	SER	DW	d.f.
TAI	3	-2.184 [2.341]	-0.070 [0.229]	-.01	9.837	0.578	54
TAI	12	-2.537 [1.786]	-0.239 [0.181]	0.07	5.829	0.153	52
AUS	3	1.443 [4.636]	-0.270 [0.473]	-.00	18.612	0.621	75
AUS	12	0.720 [2.183]	0.411 [0.264]	.07	6.435	0.311	67
SNG	3	-4.543 [1.988]	0.000 [0.353]	-.03	8.344	1.340	36
SNG	12	-4.369 [0.589]	0.050 [0.257]	-.03	2.804	0.554	32
PHL	3	6.058 [4.050	-0.259 [0.320]	-.02	17.070	1.014	37
PHL	12	12.983 [4.314]	-1.058 [0.332]	0.17	8.350	0.641	33
KOR	3	1.882 [1.298]	0.493 [0.217]	0.22	6.198	0.348	66
KOR	12	2.751 [1.230]	0.341 [0.181]	0.23	4.191	0.084	62
SAF	3	15.554 [4.227]	-1.111 [0.373]	0.18	16.459	1.227	50
SAF	12	15.375 [1.719]	-1.534 [0.128]	0.29	6.185	0.781	47
CAN	3	1.160 [1.638]	0.005 [0.457]	-.01	8.818	0.765	75
CAN	12	1.833 [2.377]	0.042 [0.598]	-.01	4.796	0.167	69

Exch Rate	Term (k)	Const.	β_1	$\bar{R}^2$	SER	DW	d.f.
ARG	3	44.730	0.628	0.18	203.529	0.354	69
		[37.446]	[0.327]				
ARG	12	55.975	0.912	0.17	149.843	0.074	51
		[74.193]	[0.682]				
MEX	3	5.194	0.065	0.04	7.107	0.446	69
		[1.636]	[0.051]				
MEX	12	4.966	0.091	0.19	4.215	0.129	51
		[1.914]	[0.046]				
BRZ	3	129.752	0.423	0.38	87.301	0.900	63
		[32.745]	[0.086]				
BRZ	12	91.164	1.061	0.61	40.601	0.431	40
		[23.321]	[0.064]				
BOL	3	26.344	-0.330	-.01	23.800	0.772	34
		[10.205]	[0.290]				
BOL	12	20.155	0.019	-.02	10.202	0.272	44
		[4.934]	[0.116]				

Notes: Term (k) is the forecast horizon in months. Figures in brackets [.] are Newey-West robust standard errors, using equal-weight windows for lag length k-1. * (**) [***] denotes significance at 10% (5%) [1%] levels.

TABLE A2
Bias in the Forward Discount
February 1988 - June 1994

$$\Delta s_{t+k} = \alpha_2 + \beta_2 fd_{t,t+k} + u_{2,t} \tag{2}$$

Exch Rate	Term (k)	Const.	ß$_2$	$\bar{R}^2$	SER	DW	d.f.
DMK	3	-1.445 [5.150]	-0.221 [1.393]	-.01	24.944	0.544	75
DMK	12	-1.939 [2.916]	0.514 [0.902]	0.01	10.205	0.240	67
FFR	3	-2.411 [6.008]	0.364 [1.207]	-.01	24.302	0.502	75
FFR	12	-5.399 [4.065]	1.733 [0.912]	0.10	10.413	0.238	67
DKR	3	-3.736 [6.881]	0.695 [1.074]	-.00	25.062	0.513	75
DKR	12	-5.793 [4.109]	1.623 [0.670]	0.09	11.028	0.239	67
UKM	3	1.712 [7.349]	0.164 [2.457]	-.01	27.674	0.608	75
UKM	12	1.515 [7.498]	0.276 [2.526]	-.01	11.688	0.254	67
NTH	3	-1.226 [5.474]	-0.236 [1.563]	-.01	25.072	0.538	75
NTH	12	-1.911 [3.060]	0.360 [0.914]	-.01	1.622	0.239	67
SFR	3	-1.199 [5.516]	0.019 [1.512]	-.01	27.480	0.528	75
SFR	12	-1.286 [3.205]	0.029 [0.828]	-.01	12.275	0.218	67

Exch Rate	Term (k)	Const.	$ß_2$	$\overline{R}^2$	SER	DW	d.f.
SKR	3	-20.890 [12.333]	4.803 [2.883]	0.17	26.640	0.563	75
SKR	12	-14.724 [6.107]	4.416 [1.373]	0.40	11.204	0.290	67
IRE	3	-4.235 [6.285]	1.273 [1.139]	0.02	24.978	0.560	75
IRE	12	-4.831 [4.437]	1.700 [0.954]	0.13	11.422	0.229	67
BFR	3	-0.060 [7.506]	-0.324 [1.625]	-.01	25.497	0.519	73
BFR	12	-4.394 [4.556]	1.379 [1.337]	0.05	11.087	0.221	65
LIR	3	-17.825 [8.107]	3.743 [1.480]	0.18	25.545	0.684	75
LIR	12	-13.773 [3.989]	3.488 [0.667]	0.39	10.417	0.373	67
NKR	3	-2.281 [5.085]	0.811 [0.368]	0.03	23.918	0.525	66
NKR	12	-5.990 [4.662]	1.923 [1.183]	0.17	9.355	0.394	62
SPN	3	-22.712 [9.039]	3.673 [1.133]	0.21	24.265	0.589	75
SPN	12	-19.438 [8.677]	3.777 [1.331]	0.35	11.433	0.310	67
YEN	3	-6.967 [4.166]	-2.905 [1.608]	0.07	20.813	0.652	75
YEN	12	-8.978 [2.367]	-4.016 [0.940]	0.59	6.121	0.515	67

Exch Rate	Term (k)	Const.	β_2	$\bar{R}^2$	SER	DW	d.f.
AUS	3	-1.199 [6.245]	0.333 [1.175]	-.01	18.677	0.621	75
AUS	12	2.993 [5.888]	-0.260 [1.025]	-.01	6.694	0.294	67
SNG	3	-6.916 [2.100]	-1.425 [0.835]	0.06	8.496	0.710	48
SNG	12	-4.598 [2.290]	0.113 [0.848]	-.02	3.214	0.296	48
SAF	3	23.519 [9.229]	-2.042 [0.923]	0.13	16.931	0.710	53
SAF	12	16.409 [3.728]	-1.437 [0.556]	0.12	6.734	0.321	69
CAN	3	6.057 [3.648]	-1.856 [0.960]	0.06	8.473	0.711	75
CAN	12	3.392 [5.535]	-0.626 [1.726]	-.00	4.768	0.158	67

Notes: SER is standard error of regression. DW is Durbin-Watson statistic. d.f. is number of degrees of freedom. Figures in the parentheses are standard errors.

TABLE A3
Test for Time Varying Risk Premium
February 1988 to June 1994

$$\Delta\hat{s}^{e}_{t,t+k} = \alpha_3 + \beta_3 fd_{t,t+k} + u_{3,t} \tag{3}$$

Exch Rate	Term (k)	Const.	ß$_3$	$\bar{R}^2$	SER	DW	d.f.
DM	3	-0.966 [1.920]	2.581 [0.420]	0.51	8.356	0.663	75
DM	12	4.476 [0.881]	1.166 [0.270]	0.45	3.595	0.686	75
FFR	3	-1.830 [2.608]	2.133 [0.645]	0.33	9.025	0.964	75
FFR	12	4.142 [1.169]	1.100 [0.303]	0.27	3.620	0.795	75
DKR	3	-3.749 [2.375]	1.939 [0.337]	0.36	8.603	0.787	75
DKR	12	3.524 [1.295]	1.049 [0.308]	0.27	3.807	0.715	75
UKM	3	5.469 [4.579]	-0.469 [0.985]	-.00	11.262	0.323	75
UKM	12	3.086 [1.888]	1.042 [0.434]	0.15	4.180	0.514	75
NTH	3	-1.637 [2.070]	2.7082 [0.468]	0.47	8.782	0.735	75
NTH	12	4.342 [0.894]	1.121 [0.264]	0.39	3.717	0.7303	75
SFR	3	1.099 [1.950]	1.871 [0.421]	0.35	9.195	0.736	75
SFR	12	5.434 [0.698]	0.910 [0.232]	0.35	3.758	1.036	75

Exch Rate	Term (k)	Const.	β_3	$\bar{R}^2$	SER	DW	d.f.
SKR	3	-6.783 [5.093]	2.013 [0.860]	0.18	11.039	0.542	75
SKR	12	0.720 [2.760]	1.347 [0.520]	0.25	4.719	0.469	75
IRE	3	-4.296 [2.571]	2.445 [0.500]	0.43	10.312	0.834	75
IRE	12	3.145 [1.334]	1.208 [0.271]	0.35	4.158	0.908	75
BFR	3	-6.853 [2.985]	3.071 [0.774]	0.46	8.538	1.124	73
BFR	12	3.706 [1.035]	1.442 [0.283]	0.37	3.591	0.843	73
LIR	3	-5.373 [3.487]	1.921 [0.487]	0.34	8.717	0.767	75
LIR	12	3.155 [1.715]	0.920 [0.257]	0.25	3.725	0.720	75
NKR	3	-5.238 [3.118]	1.404 [0.213]	0.41	11.049	0.581	66
NKR	12	0.719 [2.247]	1.267 [0.422]	0.24	5.073	0.535	66
SPN	3	-8.904 [3.958]	1.758 [0.442]	0.23	11.063	0.632	75
SPN	12	-0.389 [2.681]	1.049 [0.365]	0.18	4.970	0.693	75
YEN	3	0.641 [1.985]	1.946 [0.759]	0.18	8.866	0.796	75
YEN	12	4.915 [0.808]	1.072 [0.439]	0.18	3.912	0.618	75

Exch Rate	Term (k)	Const.	β_3	$\bar{R}^2$	SER	DW	d.f.
AUS	3	-1.450 [2.425]	1.430 [0.401]	0.30	5.478	0.854	75
AUS	12	-4.454 [1.410]	1.769 [0.292]	0.64	2.887	0.605	75
SNG	3	4.398 [0.777]	1.743 [0.370]	0.43	3.534	1.234	32
SNG	12	3.856 [0.527]	1.220 [0.309]	0.50	1.658	0.812	32
SAF	3	1.268 [3.130]	0.690 [0.357]	0.09	6.919	1.461	50
SAF	12	4.400 [1.704]	0.204 [0.235]	0.00	2.554	0.883	51
CAN	3	-1.423 [1.289]	0.750 [0.426]	0.07	3.222	0.853	75
CAN	12	-2.079 [1.332]	1.300 [0.513]	0.21	2.235	0.347	75

Notes: SER is standard error of regression. DW is Durbin-Watson statistic. d.f. is number of degrees of freedom. Figures in the parentheses are standard errors.