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TESTS OF THE FOREIGN EXCHANGE RISK PREMIUM USING THE EXPECTED SECOND MOMENTS IMPLIED BY OPTION PRICING

by

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ABSTRACT

This paper applies a new method to investigate the foreign exchange risk premium. The method is new in the sense that it utilizes the time-varying second moment expectations implied by foreign currency option pricing. The vast empirical literature on the risk premium generally neglects the role of time-varying second moments, in spite of their importance in assessing risk-return tradeoffs. In fact, this importance is borne out in the data: time-varying expectations generate valuable new evidence regarding both unbiasedness in the forward rate and portfolio balance models. Moreover, the results suggest that previous tests which assume constant second moments involve serious misspecification errors. The results also highlight the unreliability of the portfolio balance effects of sterilized intervention, in spite of the quantitative importance of expected return differentials.

Tests of the Foreign Exchange Risk Premium Using the Expected Second Moments Implied by Option Pricing

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Richard K. Lyons*

1. INTRODUCTION

The search for the foreign exchange risk premium and the variables which drive it has produced a vast empirical literature, with very limited success. However, there is reason to be cautious when interpreting many of the results, for they are generally based on an extremely tenuous assumption: the tests generally assume that conditional second moments are constant [e.g., Frankel(1982) and Hodrick and Srivastava(1984)]. Recent work by Cumby and Obstfeld(1985) and Giovannini and Jorion(1985) strongly rejects the hypothesis that the conditional variance-covariance matrix of rates of return is constant. Not only do these results suggest that many of the previous tests involve serious misspecification errors, they also have important policy implications (e.g., for the reliability of the portfolio balance effects of sterilized intervention).

This paper applies a new method for investigating the risk premium which explicitly recognizes this previously neglected dimension of the risk-return tradeoff. The approach exploits information revealed in the foreign currency options market--a market whose principal objective is the prediction of future exchange rate second moments. Using a currency option pricing model, together with realized market option prices, we solve for the market's conditional second moment expectation. Then, we use these time-varying expectations to test some hypotheses

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commonly tested in the literature to determine whether they provide any new insights.

Using the options market to measure time-varying expectations offers two key benefits. First, this measure of the second moment expectation is rooted in actual market behavior. It therefore reflects the basis on which investors are assessing risk-return tradeoffs, and is properly forward-looking. Second, it avoids the difficult question of whether to measure variance as dispersion about a conditional mean or a sample mean. Past work on asset market volatility often measures time-varying second moments using either moving variance about moving mean (MVAMM) estimators, or the ARCH model developed by Engle(1982). The option market approach differs from these in the sense that it implicitly measures the market's method of conditioning.

We first use the implied expectations to test the hypothesis of unbiasedness of forward exchange rates (which implies jointly that traders have rational expectations and that there is no foreign exchange risk premium). We do so by testing for significant correlation between expected variances/covariances and net-of-depreciation interest differentials. The results of these tests allow us to answer two key questions. First, can we reject unbiasedness using a variable upon which risk premia theoretically depend? Although the joint unbiasedness hypothesis has been strongly rejected in the literature [e. g., Hsieh(1984)], the rejections are not due to factors that theory links to the risk premium. Consequently, they provide little evidence for the alternative hypothesis that a risk premium does indeed exist. Because theory establishes a relation between second moment changes and the risk premium, this test provides a stronger basis than previous tests for interpreting rejection of the joint hypothesis as evidence for the existence of a risk premium, as opposed to a violation of rational expectations.

The second key question the unbiasedness tests will answer is: Are the Jensen's inequality terms which arise from the maximization of expected real returns in fact approximately the same magnitude as the risk premium itself? This proposition is asserted theoretically by Krugman(1981). Since these terms are functions of exchange rate second moments in his model, the proposition is easily tested within the framework of the unbiasedness tests.

We then apply some "structure" to help interpret the unbiasedness results by applying a portfolio balance model developed by Kouri(1976,1977) and Dornbusch(1983). This "intermediate" step naturally leads into a test of the Kouri-Dornbusch model itself. While this model has previously been tested by Frankel(1982), his test assumes a constant return variance-covariance matrix. Our test replicates Frankel's except for the fact that it explicitly includes a time-varying matrix. This is an important consideration; while there is a tradition in the finance literature of assuming a constant variance-covariance matrix of stock returns, the measured expectations suggest that this assumption is far less justifiable when applied to the foreign exchange market.

The paper is in six sections. Section 2 outlines the hypothesis of forward rate unbiasedness as well as the Kouri-Dornbusch model, and discusses some previous empirical results. Section 3 introduces the option pricing model used to reveal the second moment expectations. Section 4 describes the numerical procedure used to extract the implied expectations, and presents plots of the resulting series. Section 5 presents the results of the various tests. Finally, conclusions are in section 6.

2. RISK PREMIUM THEORY AND PAST EMPIRICAL RESULTS

2.1 The Hypothesis of Forward Rate Unbiasedness

The following exposition follows that in Hsieh(1984). Let s_t denote the natural logarithm of the spot exchange rate at date t, and $f_{t,1}$ the natural logarithm of the forward exchange rate contracted at date t for delivery at date t+1. $E[\, s_{t+1} \mid I(t) \,]$ is the expectation of s_{t+1} conditioned on the information set I(t), while $M[\, s_{t+1} \mid t \,]$ is the market's expectation of s_{t+1} at time t.

The hypothesis of rational expectations is that the market's expectation is the true expectation:

$$(1) \qquad M[\, s_{t+1} \mid t \,] = E[\, s_{t+1} \mid I(t) \,]$$

We will refer to this hypothesis as H1.

A second hypothesis is required to relate the forward exchange rate to expectations. The market is assumed to set the forward rate equal to the expected spot rate on delivery. That is, H2 maintains:

$$(2) \qquad f_{t,1} = M[\, s_{t+1} \mid t \,]$$

If we abstract from terms arising from Jensen's inequality,[1] sufficient conditions yielding H2 are (a) that all traders are risk neutral and (b) that markets are competitive. However, if traders are risk averse, then it is possible to obtain an equilibrium forward rate which is different from the

[1] The importance of the terms arising from Jensen's inequality relative to that of the risk premium itself is investigated in Krugman(1981). We will return to this issue in subsection 2.2.

market expectation, because of the presence of a risk premium. Thus, H2 is often called the hypothesis of 'no risk premium' in the forward rate.

Note that H1 and H2 are totally independent hypotheses. Traders may have rational expectations, but still require a risk premium for forward contracts. If they are risk averse, they may (rationally) expect a loss, in order not to have to carry any exchange risk themselves.

Also, H1 and H2 are not separately testable, because M[s_{t+1} | t] cannot be observed. However, they jointly imply:

(3) $$f_{t,1} = E[\, s_{t+1} \mid I(t)\,]$$

which is often called the hypothesis of forward rate unbiasedness in the literature.

A testable implication of (3) is the following. Define the forward rate prediction error $\varepsilon_{t+1,1}$ as:

(4) $$\varepsilon_{t+1,1} = f_{t,1} - s_{t+1}$$

Then the unbiasedness hypothesis implies that $\varepsilon_{t+1,1}$ has zero mean and is uncorrelated with any information in I(t).

Substantial empirical evidence has been built up against the unbiasedness hypothesis. Typically, researchers have interpreted rejection of this joint hypothesis as evidence that a risk premium does indeed exist. However, none of the strong rejections is due to factors that theory links to the risk premium:[2] evidence for rejection is generally either (1) serial correlation of

[2] The theoretical work bearing on the foreign exchange risk premium includes Solnik(1974), Adler and Dumas (1976), Kouri(1977), Fama and Farber(1979), Stulz(1981, 1982, 1984), Hodrick(1981), Dornbusch(1983), among others.

forward rate prediction errors[3] or (2) correlation of forward rate prediction errors with known information that is not linked theoretically to the risk premium.[4] Consequently, past rejections of unbiasedness do not offer clear evidence of a risk premium, per se.[5]

In section 5 we test whether there exists a significant correlation between expected variances/covariances and forward rate prediction errors. This test differs in an important way from previous tests because theory establishes a relation between second moment changes and the risk premium (outlined below). A rejection, then, would provide stronger evidence than previous tests for the alternative hypothesis that a risk premium exists, as opposed to the alternative of a violation of rational expectations.

2.2 A Portfolio Balance Model of the Risk Premium

The Kouri(1977) - Dornbusch(1983) model of the risk premium makes explicit a relation between the risk premium and the second moments of exchange rates. Because the derivation of the model appears elsewhere in the literature,[6] we present only a very brief description here.

Investors are assumed to maximize a function of the mean and variance of end-of-period real wealth, where bonds of different currency denominations are the only available assets. The

[3] Examples include Geweke and Feige(1979), Hansen and Hodrick(1980), Cumby and Obstfeld(1981), and Hsieh(1984).

[4] Hsieh(1984) provides the strongest evidence along these lines. His tests find a significant correlation between forward rate prediction errors and both holding period yields and forward discounts.

[5] Evidence is now appearing which points to the role of expectational errors. For example, working with survey data on exchange rate expectations, Frankel and Froot(1986) suggest that systematic prediction errors are significant in explaining forward discount bias.

[6] See Frankel(1982) for a clear exposition.

choice problem over the vector of portfolio shares generates the equilibrium relationship:[7]

(5)
$$i_t - l\, i^{\$}_t - E\Delta s_{t+1} - \mu_t = \theta\, \Omega_t\, (X_t - Q_t)$$

(6)
$$\mu_t \equiv [\, \Omega_t\, Q_t - \sigma^2_t/2\,]$$

where: $i'_t = [\ i^m_t \ \ i^p_t \ \ i^y_t\ ]$ is a vector of nominal returns on the bonds denominated in marks, pounds, and yen;[8] l is a column vector of three ones; E is the expectation operator; Δ denotes "change in"; s_t is a three-element vector of the logs of the exchange rates; μ_t are the terms arising from Jensen's inequality; θ is a coefficient of relative risk aversion; Ω_t is the variance-covariance matrix of currency depreciation; X_t is the three-element vector of asset supply shares; and Q_t is the three-element vector of consumption shares across the currencies.

Equation (5) specifies the risk premium (the right-hand-side) as a function of three different factors: the coefficient of relative risk aversion θ, the exchange rate variance-covariance matrix Ω_t, and the difference between the vector of asset supply shares and the vector of consumption shares. The formulation is based upon the case in which the consumption bundle is the same for all investors, i.e., the components of Q are the same across countries. In the case where investors in different countries have different consumption patterns, they will also have

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7 The result presented in equation (5) requires the additional assumption that goods prices are nonstochastic when expressed in the currency of the producing country. That is, the only uncertainty is exchange rate uncertainty. This assumption is made by Krugman(1981), but is considered only one special case by Dornbusch(1983). Frankel(1982) makes this assumption in his tests of the model.

8 The choice of currencies for the empirical analysis is constrained by data availability; only the mark, pound, and yen rate options provide sufficient observations.

different asset-holding preferences. An implication of this is that redistributions of wealth, for example via current account imbalances, will affect asset demands. When this change is added to the model, equation (5) takes the form:

$$(7) \qquad i_t - l\, i^{\$}_t - E\Delta s_{t+1} - \mu_t = \theta\, \Omega_t (X_t - Qw_t)$$

where Q is defined to be a matrix whose five columns indicate the consumption preferences of residents of the five countries (the fifth being the ROW), and w_t is a vector of the five wealth shares. Intuitively, the demand for a given country's asset depends positively not only on its expected relative return, but also on the wealth of those investors who have a relatively greater preference for that country's goods and thus for its assets.

Rational expectations implies that $E\Delta s_{t+1} = \Delta s_{t+1} + \varepsilon_{t+1}$, yielding the regression equation:

$$(8) \qquad i_t - l\, i^{\$}_t - \Delta s_{t+1} - \mu_t = \theta\, \Omega_t (X_t - Qw_t) + \varepsilon_{t+1}$$

This is the equation that Frankel(1982) estimates under the assumption of a constant variance-covariance matrix Ω_t. The null hypothesis of his test is that the coefficient of relative risk aversion is zero, i.e., that there exists no risk premium of the form predicted by the model. He is unable to reject the null. In section 5 we replicate Frankel's test in the sense that we include the same assets and use the same consumption shares that he does. However, this test differs in one very important way in that it does not constrain the variance-covariance matrix to be constant. This introduces a new, very volatile driving force as a complement to the much slower-moving variable $(X_t - Qw_t)$; the results should shed some light on their relative importance.

3. THE OPTION VALUATION MODEL

The basis for analysis of the value of options was provided by Black and Scholes(1973), who derived a closed-form equilibrium valuation formula for a European call option on a stock that does not pay dividends.[9] In general terms, the formula can be written as:

$$C=F(S, K, t, i, \sigma)$$

where

C=the market value of the call option
S=the price of the stock on which the option is written
K=the exercise or strike price
t=the length of time until the option expires
i=the riskless interest rate
σ=the expected standard deviation of the proportional rate of change of the stock price.

All variables appearing in this equation are directly observable except for σ, the market's collective expectation of the volatility. Thus, given the ex-post market determined value for C (together with S, K ,t, and i), one can use the model to solve for the implied volatility σ.[10]

Valuing foreign currency options requires a slightly different model because the underlying asset is an amount of foreign currency, not a share of stock. Biger and Hull(1983), Garman and Kohlhagen(1983), and Grabbe(1983) developed the model for valuing a European foreign

9 A call option on a given stock gives the owner the right to purchase one share at a predetermined price--the exercise price. A European option can be exercised only on the expiration date, while an American option can be exercised at any time up to and including the expiration date.

10 Examples of work on implied volatility include Schmalensee and Trippi(1978), Beckers(1981), and Shastri and Tandon(1986b).

exchange call:[11]

$$(9) \qquad C = S\,N(X + \sigma\sqrt{t})\,e^{-i^{*}t} - K\,N(X)\,e^{-it}$$

where

$$X \equiv \frac{\ln(S/K) + [\,i - i^{*} - (\sigma^2/2)\,]\,t}{\sigma\sqrt{t}}$$

C = the market value of the call option
S = the spot exchange rate
N( )= the cumulative normal distribution function
K = the exercise price
i = the domestic (riskless) interest rate
i*= the foreign (riskless) interest rate
t = the length of time until the option expires
σ = the expected standard deviation of the proportional rate of change of the stock price.

Unfortunately, the currency options that have been available in this country until very recently have been American options,[12] which permit exercise at any time up to and including the expiration date. One could use the European valuation formula to solve for the implied standard deviations as is done in Shastri and Tandon(1986b), but this approximation can be quite rough for call options that are likely to be exercised early. In order to generate a more precise measure from the American options, this paper uses a more general model from which the Black-Scholes model can be derived as a limiting case: the Binomial Option Pricing Method. (The model first appeared in the literature in Cox, Ross, and Rubinstein(1979); the adaptation to currency options appears in Cox(1985).) This method allows one to both (1) modify the

11 This model is identical to the one developed by Merton(1973) for equity options with the underlying security paying continuous dividends. In the case of foreign currency options, the foreign interest rate is akin to the continuous dividend rate.

12 The Chicago Board Options Exchange began trading European foreign exchange options in late 1985.

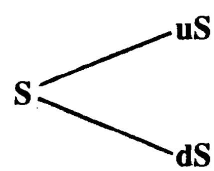
valuation procedure to take into account the different nature of the underlying asset and (2) properly include the value of early exercise. Drawing from Cox et al(1979), we begin by making three assumptions:[13]

(1) the exchange rate follows a multiplicative binomial process over discrete periods

(2) the nominal interest rate is constant and positive and

(3) there are no taxes, transaction costs, or margin requirements.

Consider the case in which only one period remains until expiration. Let S now denote the current spot exchange rate in dollars per unit of foreign currency. After one period, the exchange rate will be either uS or dS where u is the factor generated by the binomial process if the exchange rate moves upward and d the factor if it moves downward. Diagramatically:



Letting C denote the current value of a call, C_u its value at the end of the period if the exchange rate goes to uS (which represents a depreciation--i.e., more dollars to purchase a unit of foreign currency), and C_d its value at the end of the period if the spot rate goes to dS, we can diagram the call value at expiration as:[14]

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13 The final version of the model assumes that the exchange rate is distributed lognormally; this is the limiting distribution of the multiplicative binomial process we start with here.

14 A call value can never be negative, since exercise is always the option of the owner. If **uS-K** were negative, the option would expire unexercised.

$$C \begin{cases} C_u = \max[0, uS-K] \\ C_d = \max[0, dS-K] \end{cases}$$

Now, in order to value the option, it is necessary to design an equivalent (replicating) portfolio--a portfolio whose value behaves exactly as the call value does. Since by construction the call option and the equivalent portfolio will have the same value in all possible future states of nature, arbitrage will ensure that they have the same current value; consequently, the cost of the equivalent portfolio establishes the current option value. Let i denote one plus the domestic interest rate and i^* one plus the foreign interest rate. Suppose we form a portfolio with ΔS dollars invested in foreign bonds, where Δ is some number as yet undetermined. At the end of one period our investment of ΔS dollars will take on one of two values <u>in dollars</u>:

$$\Delta S \begin{cases} u\Delta S i^* \\ d\Delta S i^* \end{cases}$$

That is, both the binomial process and the foreign interest rate are affecting its dollar value.

In order to build a replicating portfolio, it is necessary to combine this with an investment of B dollars (as yet undetermined) in domestic bonds.[15] The end of period dollar value of this portion of the portfolio will necessarily be Bi. Thus, at the end of one period the whole portfolio

[15] The intuition behind this is that the dollar value of the portion of our replicating portfolio that is invested in foreign bonds will be affected one-for-one by the exchange rate change. The value of a call option on foreign exchange, on the other hand, will not be affected one-for-one by an exchange rate change, but rather by some fraction of the change. If we let this fraction be denoted by Δ, then the proportion of our replicating portfolio that is invested in domestic bonds (the value of which is unaffectecd by exchange rate movement) will be $1-\Delta$, exactly matching the call option's degree of insensitivity to spot changes.

will take on one of the two following dollar values:

$$\begin{tikzcd}[row sep=small] & \Delta uSi^*+Bi \\ \Delta S+B \arrow[ur] \arrow[dr] & \\ & \Delta dSi^*+Bi \end{tikzcd}$$

Since Δ and B can be set to any value one chooses, one can choose those values such that the two possible end-of-period portfolio values exactly match the end-of-period call values C_u and C_d. This provides an equivalent portfolio, and if there are to be no riskless arbitrage opportunities, the current value of the call, C, must equal the cost of the equivalent portfolio. Setting these values equal to each other yields:

(10) $$\Delta uSi^*+Bi=C_u$$

(11) $$\Delta dSi^*+Bi=C_d$$

Equations (10) and (11) allow us to solve for the two values for Δ and B that generate the replicating portfolio. After a little algebra:

$$\Delta = \frac{Cu - Cd}{S\, i^*\, (u - d)} \qquad \text{and} \qquad B = \frac{uCd - dCu}{i\, (u - d)}$$

Since the value of the call must be the same as the value of the equivalent portfolio, we can simply plug in these values for Δ and B :

(12) $$C = \Delta S + B$$

(13) $$= \frac{(C_u - C_d)\,S}{S\,i^*\,(u-d)} + \frac{uC_d - dC_u}{i\,(u-d)}$$

(14) $$= \frac{\left(\frac{i}{i^*} - d\right) C_u + \left(u - \frac{i}{i^*}\right) C_d}{i\,(u-d)}$$

Equation (14) establishes the one period pricing relation given values for u, d, i, i*, C_u, and C_d, all of which are known at the current time.[16] If we now consider a case in which the number of periods to expiration (henceforth denoted n) is greater than one, the logic behind the valuation of a multiperiod call begins to assert itself. For the case of n=2 we have:

$$\begin{tikzcd}[column sep=small, row sep=small]
 & & C_{uu}=\max[0,uuS-K] \\
 & C_u \arrow[ur] \arrow[dr] & \\
C \arrow[ur] \arrow[dr] & & C_{du}=\max[0,duS-K] \\
 & C_d \arrow[ur] \arrow[dr] & \\
 & & C_{dd}=\max[0,ddS-K]
\end{tikzcd}$$

The values of u ,d ,S, and K (known at the current time) establish the three possible call values of C_{uu}, C_{du}, and C_{dd} that could be realized at the date of expiration.[17] This is the starting point in

16 It can be shown that the choice of $u=\exp(\sigma\sqrt{t/n})$ and $d=1/u$ provide the final model with the proper limiting properties, where σ is the (annualized) standard deviation of the proportional changes in the spot rate, t is the time to expiration in years, and n is the number of subintervals used in the approximation. See Cox and Rubinstein (1985) pg. 242.

17 At this point we are still assuming that the option cannot be exercised until the expiration date.

the valuation process. We can work leftward through the tree to solve for values of C_u and C_d given our pricing relation in equation (14) (together with values of i and i*). Once C_u and C_d are established, one more step leftward using the pricing relation establishes the current call value, C. Thus, although this is not a closed form solution as in the Black-Scholes case, in general terms we have the formula:

$$C = F(S, K, t, i, i/i^*, \sigma, n) \quad (14)'$$

If the assumptions underlying the model are correct, as the number of subintervals n approaches infinity, the call value calculated should converge to the true value.[18]

The benefit of the binomial procedure is that it allows, at each node of the tree, for a comparison of the value if exercised immediately with the value if kept alive (unexercised). At those nodes where immediate exercise is optimal, the exercise value is substituted for the value calculated using the pricing relation in equation (14). This ensures that full consideration is given to the impact of the early exercise feature of an American option.[19]

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18 The assumptions underlying the model are: (1) the spot rate is log-normally distributed with constant variance, (2) the risk-free interest rates are known and constant, and (3) capital markets are frictionless.

19 On a stock that pays no dividends, the values of American and European call options are identical; since no "rents" (dividends) are being forgone by not exercising an American call, it will always be worth more alive than exercised. However, in the case of foreign exchange options, rents are being forgone: the interest rate applying to the foreign currency if the option were exercised. Thus, it's the differentials between domestic and foreign interest rates that make the optimality of early exercise possible.

4. THE IMPLIED SECOND MOMENT EXPECTATIONS

4.1 Inputs For The Valuation Model

American spot currency options began trading on the Philadelphia Exchange (PHLX) in early 1983. Beginning in September 1985, the Chicago Board Options Exchange (CBOE) began trading European currency options. In order to generate a sufficiently long time series, we use the PHLX option prices to recover the second moment expectations. Additionally, while the PHLX now trades options on six major currencies, the data are too sparse on all but the three most heavily traded currency options: the British Pound, the West German Mark, and the Japanese Yen.

A very important consideration when attempting to back implied variances out of an option pricing model is the synchronous data problem. A spot price that does not correspond to the timing of the option trade will distort the variances if the spot rate differs from the rate prevailing at the time of the trade. For this reason, we use a transactions data base which provides the underlying spot rate at the time of each option trade (the PHLX-OSU Currency Options Data Base, compiled with the support of the Philadelphia Stock Exchange by James N. Bodurtha Jr.). The data base runs from February 28th, 1983 through June 27, 1985.

The weekly data used in the simple efficiency regressions covers a sample period from July, 1983 through May, 1986, for a total of 151 observations (Wednesday closing prices). The first four months of the transactions data are not included due to the thinness of trading; there were eight trading days between March and June, 1983, on which no call option traded for at least one of the three currencies, while this occurred only once after June, 1983. (The closing trade of the previous day is used). In order to lengthen the sample beyond June 27, 1985 we use the Philadelphia Exchange closing option and spot prices as reported in the New York Times (46 of the 151 observations). The spot rate that is printed in the Times is the Philadelphia Exchange spot price at the close of the options market (2:30 Eastern Time). Since trading volume is quite high through the later third of the sample, the 2:30 spot price is likely to be a very good approximation to the spot rate prevailing at the time of the final option trade. The resulting series do not appear to be affected by the change.

For each observation we choose the call option that is closest to being at-the-money, with a time to maturity between three and six months. There are a number of justifications for focusing on the at-the-money options.[20] First, the Black-Scholes model has a tendency to work best for options nearest the money. Second, the partial derivative of the call price with respect to the future volatility is higher the closer the call is to being at-the-money, suggesting that the prices of these options are more sensitively linked to the underlying volatility. And third, these options are more heavily traded. The choice of the time to maturity window between three and six months is motivated by a balancing of the increasing significance of the method's approximations at shorter maturities against the increasing thinness of trading at longer maturities.

Recall that the Binomial Currency Option Pricing Model establishes a foreign exchange option's value as a numerical approximation using seven variables:

$$C = F(S, K, T, i, i / i^*, \sigma, n) \qquad (14)'$$

For each implied volatility calculation, S is taken to be the spot price of the corresponding currency from the transactions data base (PHLX spot rate at option market close after July, 1985), K to be the exercise price of the option, t to be the number of calender days until the last trading day prior to expiration, and the number of subintervals n to be 50 (see subsection 4.2 for more on the significance of n).

20 Stan Beckers(1981) has demonstrated that, with implicit variances computed from daily closing option and stock prices, using only the ~~implicit variance~~ of the option nearest the money produces as good a prediction of future variance as other more elaborate selection/weighting schemes.

The relevant interest rates are obtained as follows. The domestic interest rate is calculated using the rates implied by the average of the Bid and Ask prices for the U. S. Treasury Bill expiring on the date closest to the expiration of the option. The ratio i/i* is obtained using the currency futures contract expiring at the same time as the option, together with the covered interest parity relation.[21] The Treasury Bill discounts and futures prices are from the New York Times (Wednesday closing prices).

4.2 Generating The Expected Variance Series

The binomial valuation model of section 3 requires working through a binomial "valuation tree" just like the one presented there for the case of n=2 subintervals. In calculating the final implicit volatilities, we set n equal to 50 (a number of subintervals that provides a "sufficiently" close approximation to the Black-Scholes value when valuing a stock call option).[22]

The numerical approximation procedure used starts with an initial value for the implicit volatility and combines it with the other determining variables in the model. After going through the tree completely, the procedure generates a call value. This call value is then compared to the actual market price. At this point an adjustment is made to the initial volatility value according to the Newton-Raphson numerical search procedure.[23] The process then iterates until a change in the annual spot volatility (standard deviation) of less than .0001 is encountered.[24]

The resulting three series of the expected variance (of changes in the log of the exchange

[21] Violations of covered interest parity represent riskless arbitrage opportunities; the condition holds very well across assets within the same jurisdiction. For evidence see Frenkel and Levich(1979).

[22] See Cox and Rubinstein, Options Markets, p. 243, for examples of the accuracy of the Binomial Approximation to the Black-Scholes values for different values of n.

[23] See Manaster and Koehler(1982) for further information regarding the search procedure.

[24] Since the derivative of a call value with respect to the volatility is always positive, if an implied volatility exists then it is unique.

rate on an annual basis) are plotted on the following page together with a measure of the historical mean squared change in the log of the exchange rate calculated from the previous month's daily spot movements (20 business days) according to:

$$(15) \qquad RVAR_t = \sum_{j=t-19}^{t-1} (R_j - \bar{R})^2 / 18$$

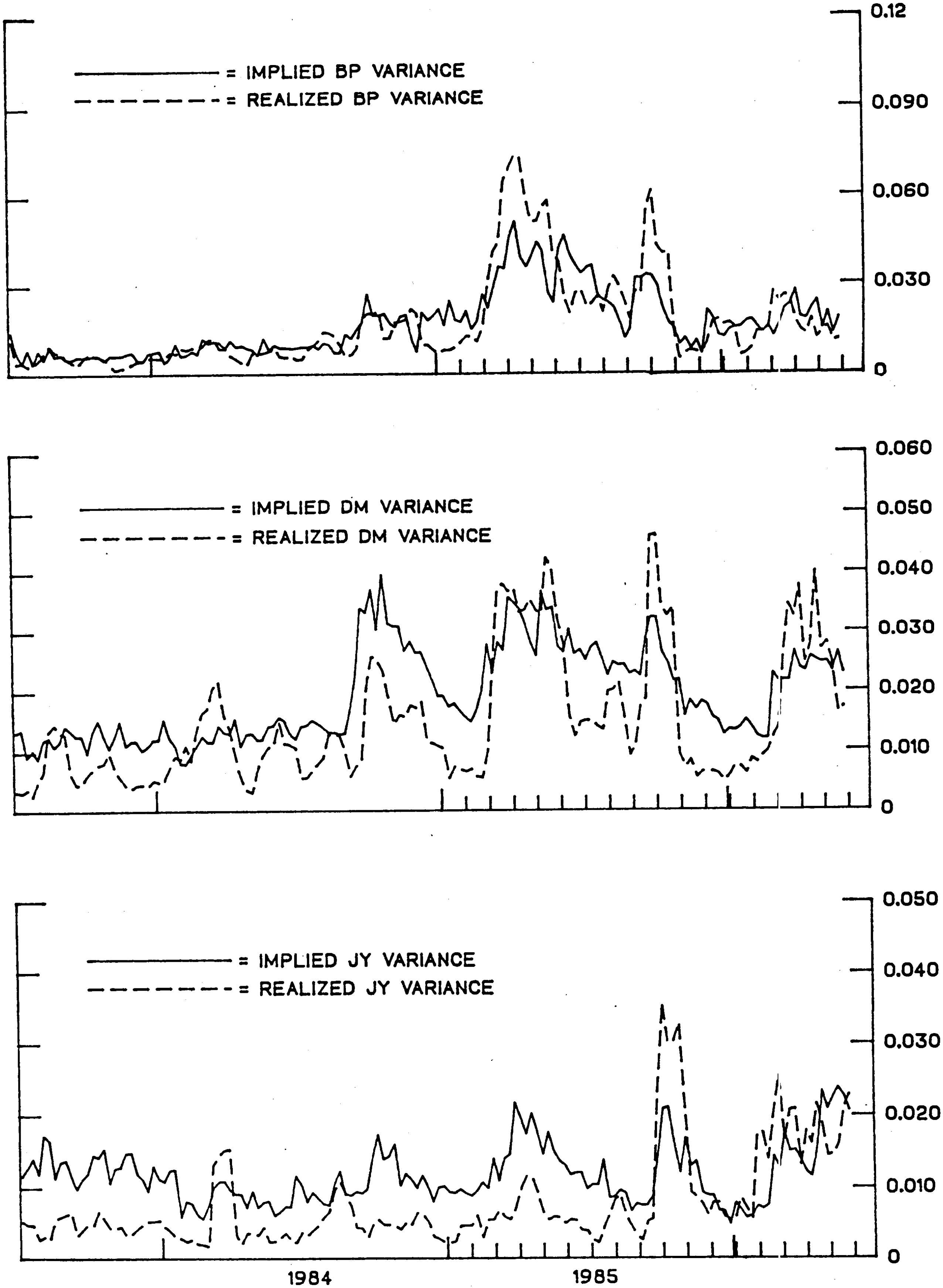
where $$R_j = \ln(S_j) - \ln(S_{j-1})$$

and $$\bar{R} = \sum_{j=t-19}^{t-1} R_j / 19$$

Since $RVAR_t$ in (15) is a measure of the variance of daily changes in the log of the exchange rate, we use an assumption of independence to arrive at the variance on an annual basis by multiplying by 262. This is the series that appears as the REALIZED VARIANCE in the plots.[25]

25 The horizon over which the REALIZED VARIANCE is calculated is chosen so as to generate similar volatility in the two series; a longer horizon smooths the series while a shorter one increases the volatility. Econometric results bearing on the question of which horizon best "explains" the implied variance series are not included in this paper.

VARIANCES OF ANNUAL LOG SPOT RATE CHANGES



5. EMPIRICAL RESULTS

Abstracting from terms arising from Jensen's inequality, forward rate unbiasedness implies that forward rate prediction errors have zero mean and are uncorrelated with any information available to agents at the time the forward rate is determined. In subsection 2.1, we defined the one-period forward rate prediction error as $\varepsilon_{t+1,1} \equiv f_{t,1} - s_{t+1}$ where $f_{t,1}$ is the log of the one-period forward rate set at time t and s_{t+1} is the log of the realized spot rate at $t+1$. As is common in the literature [e.g., Hsieh(1984)], we assume that covered interest parity holds, which implies that:

$$(16) \qquad i^*_t - i_t - \Delta s_{t+1} = \varepsilon_{t+1}$$

where i^*_t and i_t are the foreign and domestic nominal interest rates, Δs_{t+1} is equal to the change in the log of the spot rate between t and $t+1$, and ε_{t+1} is the prediction error realized at $t+1$ (we will suppress the one-period subscript from now on). Notice that the left-hand-side is the same as that which appears in the Kouri-Dornbusch model. That is, the test of the Kouri-Dornbusch model is really a test of forward rate unbiasedness in which the exact form of the bias--the risk premium--is specified. This fact will be helpful for interpreting the coefficients in the unbiasedness tests.

First, we test to see if expected exchange rate variances and covariances are correlated with the prediction errors. In light of the Krugman(1981) results, however, the prediction errors should not be orthogonal to these variables due to the Jensen's inequality terms. Under the null hypothesis of unbiasedness, the prediction errors netted of these terms are orthogonal to all information. For example, in the case of the pound equation we have:

$$(17) \qquad i^p_t - i_t - \Delta s_{t+1} - \mu^p_t = \varepsilon_{t+1}$$

where $$\mu^p_t \equiv [\,(Q_p - 0.5)\,\sigma^2_{p,t} + Q_m\,\sigma_{pm,t} + Q_y\,\sigma_{py,t}\,]$$

as per equation (8) with no risk premium. The OLS results for the single equation unbiasedness test appear in Table 1 below (weekly data). The regressors are the implied variances and covariances as described below the table.

TABLE 1

$$i^*_t - i_t - \Delta s_{t+1} - \mu_t = \beta_0 + \beta_1{}^*X_1 + \beta_2{}^*X_2 + \beta_3{}^*X_3 + \varepsilon_{t+1}$$

	ß0	ß1	ß2	ß3	F	R2	DW
MARK	-.001 (0.20)	-16.9 (0.49)	-54.1 (1.89)	133 (1.91)	.03	.07	1.77
POUND	.002 (0.39)	19.6 (0.28)	-62.7 (1.93)	139 (1.05)	.03	.08	2.05
YEN	.000 (0.06)	-69.3 (1.19)	-20.1 (0.40)	49.8 (1.80)	.17	.03	1.46

* T-statistics using White matrix in parentheses. "F" denotes marginal significance level of the test that $\beta_1=\beta_2=\beta_3=0$.

Mark equation:	X1 = variance of log changes in the dollar/mark rate (weekly basis) X2 = covariance of log changes in the mark rate and pound rate X3 = covariance of log changes in the mark rate and yen rate
Pound equation:	X1 = covariance of log changes in the pound rate and mark rate X2 = variance of log changes in the pound rate X3 = covariance of log changes in the pound rate and yen rate
Yen equation:	X1 = covariance of log changes in the yen rate and mark rate X2 = covariance of log changes in the yen rate and pound rate X3 = variance of log changes in the yen rate

These regressions are run on the full sample period from July, 1983 through May, 1986 (151 weekly observations). The interest rates are 7-day Eurocurrency deposit rates expressed on a weekly basis. We generate the covariance terms by assuming that the correlations are constant, and then using the relation $\mathrm{Cov}(X,Y) = \rho_{X,Y}\,\sigma_X\,\sigma_Y$ where ρ is taken to be the sample correlations, and the standard deviations are the square roots of the implied variances.[26]

The results provide important evidence on two fronts. First, the results of the F-tests, together with the results for individual coefficients, provide substantial evidence against the joint unbiasedness hypothesis, particularly in the cases of the mark and pound. While this evidence for rejection is not as strong as that which appears elsewhere in the literature, it is qualitatively very different in that it is due to variables which theory links to the risk premium. In this sense, it is better evidence for the alternative hypothesis that a risk premium does indeed exist, as opposed to the alternative of a violation of rational expectations.

The second point to be gleaned from Table 1 concerns the relative magnitude of the Jensen's inequality terms and the risk premium. Krugman(1981) demonstrates that in theory they are approximately the same magnitude. However, the theoretical coefficients on X_1, X_2, and X_3 arising solely from Jensen's inequality are all fractions less than one, per equation (17).

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26 The sample correlations are: $\rho_{pm} = 0.736$, $\rho_{py} = 0.378$, and $\rho_{my} = 0.550$.

The results in Table 1 demonstrate that the coefficients β_1, β_2, and β_3 (which are net of the theoretical effects from Jensen's inequality due to subtracting μ_t from the left-hand-side) are generally much larger than the fractional coefficients imputed to Jensen's inequality.

Consider now an interpretation of the coefficients which derives from the portfolio balance model of Kouri and Dornbusch. The model, described by equation (8), specifies the risk premium (RP) on currency i as:

$$RP_{i,t} = \theta\,\Omega_{i,t}\,(X_t - Q_t) = \theta(X_{i,t} - Q_{i,t})\sigma^2_{i,t} + \theta(X_{j,t} - Q_{j,t})\sigma_{ij,t} + \theta(X_{k,t} - Q_{k,t})\sigma_{ik,t}$$

where θ is the coefficient of relative risk aversion, Ω_i is the i^{th} row of the variance-covariance matrix of exchange rate changes, X_t is the vector of asset supply shares denominated in the different currencies, and Q_t is the vector of consumption shares in the different currencies. If one assumes that the difference between the asset supply shares and the consumption shares $(X_t - Q_t)$ is constant, then each of the coefficients in Table 1 should be the same across the three equations. That is, $\beta_1 = \theta\,(X_p - Q_p)$, $\beta_2 = \theta\,(X_m - Q_m)$, and $\beta_3 = \theta\,(X_y - Q_y)$. Results of OLS estimation of the pooled regression are reported in Table 2 below.

TABLE 2

$$i^*_t - i_t - \Delta s_{t+1} - \mu_t = \beta_0 + \beta_1{*}X_1 + \beta_2{*}X_2 + \beta_3{*}X_3 + \varepsilon_{t+1}$$

	β_0	β_1	β_2	β_3	R2	DW
POOLED	.005 (2.24)	9.92 (0.99)	-22.6 (2.70)	-19.2 (1.69)	.03	1.84

* T-statistics using White matrix in parentheses.

The significance of the coefficient β_2 is even stronger evidence against the null hypothesis that all information is orthogonal to the left-hand-side. Additionally, assuming risk averse agents, the negative sign of β_2 suggests that investors perceive the existing share of pound-denominated assets to be lower than that share which minimizes consumption risk. The insignificance of β_1 and β_3 is consistent with the perception that asset shares for the mark and yen are more closely in balance with the minimum risk shares. The implication for θ of the magnitudes of the coefficients is puzzling, however. Since (X - Q) is the difference of two fractions (shares), the magnitude of β_2 suggests that investors are "extremely" averse to exchange rate risk, i.e that θ is <u>much</u> larger than is commonly believed.

The analysis above naturally leads us to a test of the Kouri-Dornbusch model itself. Once we have a measure of the asset supply shares and the consumption shares, we can estimate the single coefficient θ in (8):

$$(8) \qquad i_t - l\, i^{\$}_t - \Delta s_{t+1} - \mu_t = \theta\, \Omega_t\, (X_t - Qw_t) + \varepsilon_{t+1}$$

As previous results point out, however, given the correlations of the variables in Ω_t with the left-hand-side, the estimated coefficient θ will be very sensitive both in sign and magnitude to the sign and magnitude of the relatively stable $(X_t - Qw_t)$. This said, we choose to measure the asset shares as is done previously in the literature by Frankel(1982), so that the effect of changing second moments can more easily be discerned. The world portfolio which he considers includes only outside government debt.[27]

[27] See the Data Appendix for definitions.

Frankel estimates θ under the strong assumption of a constant variance-covariance matrix Ω_t. The null hypothesis of the test is that the coefficient of relative risk aversion is zero, i.e., that there is no portfolio balance risk premium of the form the model predicts. He is unable to reject. The test below also tests for the null hypothesis of a zero coefficient of risk aversion. However, this test permits the covariance matrix to vary. The elements of each month's covariance matrix are determined by the implied expected variances from the options market (monthly basis), with the covariances determined as above using the assumption of constant correlations.

The results of the OLS regressions are presented in Table 3. They cover monthly observations from February, 1983 through December, 1985, for a total of 35 months. The EQUAL CSN SHARES regression corresponds to the case in which all countries consume the same basket of goods (see appendix); the interest rates are those for one month Eurocurrency deposits (monthly basis); the spot rate is that used to generate the implied expected variances; the coefficient β_1 corresponds to the coefficient of relative risk aversion:

TABLE 3

$$i^*_t - i_t - \Delta s_{t+1} - \mu_t = \beta_0 + \beta_1 * \Omega_t (X_t - Qw_t) + \varepsilon_{t+1}$$

	ß0	ß1	R2	DW
EQUAL CSN SHARES	-.006 (1.38)	-241 (2.74)	.07	2.09
CSN SHARES VARY	.007 (1.22)	-172 (1.49)	.02	2.10

* T-statistics in parentheses

The results are both discouraging and interesting. The coefficient of relative risk aversion (β_1) is both negative and much larger in absolute value than one would expect. Thus, the test results provide no support for the model. The existence of a "pole" at the point $(X_t - Qw_t) = 0$ is a possible explanation of the wrong sign, however. That is, given the demonstrated relationship between the relatively volatile Ω_t and the left-hand-side, measuring the relatively stable $(X_t - Q_t)$ with a sign that is opposite that perceived by the market would generate a negative coefficient rather than a positive one if the model were in fact true. The extremely significant coefficient in the equal-consumption-shares case suggests that a relationship exists. At the very least, it constitutes a very strong rejection of the unbiasedness hypothesis.

In order to gauge the sensitivity of the results to the consumption shares, we examine two alternative cases: (1) investors measure consumption risk against dollar-denominated consumption, so that the dollar share in consumption is one and (2) investors measure consumption risk wholly against home-denominated consumption, so that the home share equals one. Table 4 presents the results of the OLS regressions:

TABLE 4

$$i^*_t - i_t - \Delta s_{t+1} - \mu_t = \beta_0 + \beta_1 {*} \Omega_t (X_t - Q_t) + \varepsilon_{t+1}$$

	β_0	β_1	R2	DW
DOLLAR SHARE = 1	-.015 (1.69)	-44.6 (1.87)	.03	2.19
HOME SHARE = 1	.009 (1.23)	-209 (1.44)	.02	2.10

* T-statistics in parentheses

In the dollar-share-equals-one regression the coefficient of relative risk aversion now has the correct sign. While this case is less plausible than the others, the result underscores the sensitivity to the measurement of the difference (X_t - Q_t).

As a final measure to help identify the role of various factors in producing the estimates above, we regress the left-hand-side on the asset supply share series alone in bilateral equations. The theory predicts that the coefficients on asset supplies should be positive: a higher asset supply requires a higher expected return to induce investors to hold them. Table 5 presents the results of the OLS regressions:

TABLE 5

$$i^*_t - i_t - \Delta s_{t+1} - \mu_t = \beta_0 + \beta_1{}^*(\text{ASSET SHARE})_t + \varepsilon_{t+1}$$

	ß0	ß1	R2	DW	DF
MARK	.083 (1.96)	-1.37 (2.05)	.11	2.30	33
POUND	.096 (1.34)	-0.94 (1.35)	.05	2.09	33
YEN	.324 (3.02)	-1.39 (2.99)	.21	1.96	33

* T-statistics in parentheses.

All of the coefficients on the asset supply shares are of the wrong sign, significantly so in the case of the yen and mark. To the extent that the measured shares represent the relevant assets, this is further evidence against the theory. Additionally, these results shed some light on why the measured coefficients of risk aversion are typically negative.

In addition to the above analysis, the volatility of expectations in itself has potential policy implications, e.g., for the reliability of the portfolio balance effects of sterilized intervention. Consider for illustration a two-country model of the risk premium:

$$i^*_t - i_t - E\Delta s_{t+1} = \theta \sigma^2_t (X_t - Qw_t) \tag{23}$$

A common view is that sterilized intervention (instrument) affects the exchange rate (target) according to the above relationship with a stable, if not constant, exchange rate variance. However, the pronounced volatility of market's second moment expectation seriously undermines this view. The extreme noise introduced would swamp any policy induced variation in X_t due to sterilized intervention, with the sign of the effects of changes in σ^2_t depending on the sign of $(X_t - Qw_t)$. Additionally, to the extent that current account imbalances affect exchange rates through the above relationship (by reallocating w), their role as a driving force pales in comparison to the role of σ^2_t.

6. CONCLUSIONS

The foreign currency option market provides a unique method for measuring ex-ante variance expectations: beliefs embedded in option prices are rooted in actual market behavior and reflect the "true" information set available to investors. Moreover, the model used to recover the implied expectations is preference-free, relying wholly upon arbitrage pricing methods.[28] The result is a series which corresponds quite closely to actual variances, and provides some insight into how far back investors look in forming their expectations.

Overall, the results of the tests support one conclusion above all others: changes in the market's second moment expectations are systematically related to expected return differentials. The joint unbiasedness hypothesis is rejected in a number of different tests. Rejection of this joint hypothesis is not new, however. What is new is that the rejections are due to variables which theory ties to the risk premium. In this sense they provide much better evidence in favor of the alternative hypothesis that a risk premium does indeed exist, as opposed to the alternative of a violation of rational expectations.

The test of the Kouri-Dornbusch model of the risk premium finds little support in the data. The absolute size of the measured coefficients of relative risk aversion are broadly consistent with the magnitudes of the coefficients in the unbiasedness regressions, but the sign is typically wrong. There are at least two explanations for this result which are consistent with the validity of the theory. First, as explained in section 5, the sign and magnitude of the measured coefficient are very sensitive to the asset and consumption share definitions, given the correlation between the elements of the variance-covariance matrix and the dependent variable (as demonstrated in the unbiasedness tests). This suggests the possibility of experimenting with broader asset definitions than those used in Frankel(1982). A second explanation works from the other side of the equa-

[28] This method of measuring ex-ante investor uncertainty lends itself to a number of different applications. The author is currently using the method to investigate time-varying risk premia in the term structure of interest rates.

tion. That is, perhaps replacing expected depreciation with realized depreciation plus a random error term is off the mark, especially during the sample period considered (1983 - 1986). Working with survey data on exchange rate expectations, Frankel and Froot(1986) find evidence that systematic prediction errors are playing a significant role. A possible remedy for this would be to replace the realized depreciation with the median survey expectation on the left-hand-side. The results in this case are likely to be quite different from the current results in light of the fact that most all of the variability in the dependent variable is due to this component.

A final note concerns the more general implications of rapidly varying second moment expectations. First, they call into question the reliability of the portfolio balance effects of sterilized intervention. Feasible policy-induced changes in relative asset supplies are dwarfed by autonomous movements in investors' second moment expectations. And second, theoretical and empirical work which neglects the perceived nonstationarity of exchange rate second moments is not consistent with the actual marketplace.

DATA APPENDIX

Rates of Return

The rates of return used for the regressions reported are the seven-day and one month Eurocurrency interest rates as reported in the London Financial Times. Wednesday closing rates are used for the weekly data, while rates for the last business day of the month are used for the monthly data. The depreciation of the spot rate is measured as the change in the logarithm of the spot rate that is used to calculate the implied option volatility each week/month (Philadelphia Exchange). This corresponds to the transactions data described in the text from July, 1983 through June, 1985, and to the Philadelphia option market closing spot reported in the New York Times for July, 1985 through May, 1986.

U.S. Dollar Assets

DOASST = world supply of dollar assets in billions of dollars, end of month. Calculated as DODEBT + FEDINT - NDOLCB.

DODEBT = gross public debt of the U.S. Treasury and other U.S. government agencies, excluding that held by U.S. government agencies and trust funds, i.e. held by the Federal Reserve, private domestic investors, and foreigners, at end of month (source: Treasury Bulletin, Table FD-1); minus two issues of 'Carter Bonds,' which are denominated in foreign currency: $1595.2 million dating from December 1978 and another $1,351.5 million from March 1979 (source: Federal Reserve press release, June 1979).

FEDINT = dollars supplied by the Fed in cumulative foreign exchange intervention. computed by $FEDINT_t = FEDINT_{t-1} + \Delta FEDINT_t$, on a benchmark of the dollar value of all U.S. international reserve assets (gold, foreign exchange, SDRs, and IMF position) in February, 1983 (source: Federal Reserve Bulletin, Table 3.12).

ΔFEDINT = intervention, equal to increases in reserves, corrected for valuation changes. Computed as the change in gold holdings (there were no valuation changes during the sample period), plus the change in foreign exchange holdings in dollars minus valuation change (last period's foreign exchange times the change in the dollar/mark rate; most of the holdings have been in marks), plus change in SDRs and IMF position in dollars minus valuation change (last period's SDRs and IMF position times the change in the dollar/SDR rate). (Source for reserve holding: Federal Reserve Bulletin, Table 3.12. Source for dollar/SDR rates: IMF Internationl Financial Statistics, line sa).

NDOLCB = holding of dollar assets (regardless whether government securities) by foreign central banks as foreign exchange reserves. Millions of dollars, end of month. These data are available on an annual basis from the IMF Annual Report, 1985 and 1986, Table 15. The numbers are multiplied by the dollar/SDR rate to get the dollar value of official holdings of U.S. dollars at the end of 1983:12, 1984:12, and 1985:12. The monthly numbers are arrived at by linear interpolation.

Deutsche Mark Assets

DMASST = world supply of mark assets in millions of marks, end of month. Calculated as DMDEBT + BBINT - NDMCB.

DMDEBT = debt of the German Federal Government, end of month. Source: Monthly Report, Deutsche Bundesbank, Table VII 9.

BBINT = cumulative Bundesbank sales of mark assets for international reserves in exchange market intervention, calculated as GRES - GADJ.

GRES = net external position of the Bundesbank, valued in millions of marks, end of month. Source: Monthly Report, Deutsche Bundesbank, Table IX 6(a).

GADJ = 'balancing item' to the Bundesbank's external position, and adjustment by the Bundesbank every December to reflect capital gains on foreign exchange and other reserves. These items must be taken back out of GRES so that only changes in reserves due to purchases or sales of mark assets are counted. Cumulated with a benchmark of zero in 1983:2. Source: Monthly Report, Table IX, col. 7.

NDMCB = holdings of mark assets (regardless whether government securities) by foreign central banks as foreign exchange reserves. These data are available on an annual basis from the IMF Annual Report, 1985 and 1986, Table 15. The numbers are converted to marks by using the dollar/SDR exchange rate (from IFS line sa) and the mark/dollar exchange rate (from IFS line ae). Monthly numbers are obtained by linear interpolation.

Pound Sterling Assets

PSASST = world supply of pound assets in millions of pounds, end of month. Calculated as PSDEBT + BEINT - NPSCB.

PSDEBT = pound sterling debt of the British government. Computed as the cumulation of the public sector borrowing requirement (Central Statistical Office Financial Statistics, Table 2.5, col. 14) on a 1983:3 benchmark of 127,727 million pounds (CSO Financial Statistics, Table S4).

BEINT = cumulative Bank of England sales of pound assets for international reserves in exchange market intervention. Computed by $BEINT_t = BEINT_{t-1} + \Delta BEINT_t$ (U.K. Balance for Official Financing; source: CSO Financial Statistics, Table 10.1, divided by three to get monthly numbers), on a 1983:2 benchmark of total reserves minus gold in dollars (from IMF IFS, line 11.d) times the pound/dollar exchange rate (from IFS line ag).

NCSCB = holdings of pound assets (regardless whether government securities) by foreign central banks as foreign exchange reserves. These data are available on an annual basis from the IMF Annual Report, 1985 and 1986, Table 15. The numbers are converted into pounds by using the pound/dollar exchange rate (from the IFS, line ae) and the dollar/SDR exchange rate. Monthly numbers are obtained by linear interpolation.

Japanese Yen Assets

JYASST = world supply of yen-denominated assets in millions of yen, end of month. Calculated as JYDEBT + BJINT - NJYCB.

JYDEBT = yen-denominated debt of the Japanese government not held by the government itself. Computed as the Total Debt minus Total Government Holdings in millions of yen (source: The Bank of Japan Economic Statistics Monthly, Table 82).

BJINT = cumulative Bank of Japan sales of yen assets for international reserves in exchange market intervention. Computed by (1) the change in foreign exchange reserves corrected for valuation changes, equal to the yen/dollar exchange rate (from IFS, line ae) times the change in foreign exchange reserves expressed in dollars (from IFS line 1.dd) under the assumption that most foreign exchange reserves are held as dollars; plus (2) the change in the IMF position and holdings of SDRs corrected for valuation changes, equal to the yen/dollar exchange rate times the difference between this period's IMF position and SDR holding (from IFS, lines 1.cd and 1.bd, respectively) and last period's, multiplied buy the new over old dollar/SDR exchange rate; cumulated on a February 1983 benchmark of (3) the yen value of total international reserves, equal to the yen/dollar exchange rate times 23,930 dollars (from IFS, line 1). There were no changes in gold stocks over the sample period.

NJYCB = holdings of yen assets (regardless whether government securities) by foreign central banks as foreign exchange reserves. These data are available on an annual basis from the IMF Annual Report, 1985 and 1986, Table 15. The numbers are converted to yen by using the yen/dollar exchange rate (from IFS, line ae) and the dollar/SDR exchange rate. Monthly numbers are obtained by linear interpolation.

Wealth

The net wealth of the citizens of each of the four countries, in millions of local currency units, end of month, is calculated as the cumulation of the government deficit plus the current account surplus. The government deficit for the U.S. is the change in the gross public debt for the U.S. Treasury, excluding only that held by U.S. government agencies and trust funds (DODEBT not corrected for the two issues of 'Carter Bonds'); for Germany and Japan it is the change in government debt (DMDEBT and JYDEBT respectively); for the U.K. it is the Public Sector Borrowing Requirement.

The monthly current account for each country is obtained by taking the current account for a given quarter, subtracting the balance of trade for that quarter to obtain the balance on services and transfers, dividing by three to get a monthly estimate, and adding to the balance of trade, which is available monthly. The quarterly current accounts for all four countries come from IMF IFS, line 77a.d. The monthly balance of merchandise trade for each country is taken as the difference between exports and imports (IFS, lines 70 and 71).

The benchmarks for the accumulation of wealth are at year end 1972. They are very ad hoc, since accurate data on the level of wealth are impossible to get, and the benchmark is essentially only a constant term in the equation anyway. For the U.S. the wealth benchmark is 400 billion dollars, taken from Dooley and Isard (1979, p. 24), and 'estimated from end-of-1972 stocks in Federal debt, monetary base, and net claims on foreigners. For Germany, the U.K., and Japan the wealth benchmarks are taken as the ratio of the dollar value of own GNP (from IFS) to U.S. GNP in 1983, times U.S. wealth at 1983:2 (which includes the cumulation of U.S. current accounts and deficits between 1972:12 and 1983:2).

Consumption Shares

When residents of all countries are assumed to consume the same basket of goods, the share of world consumption allocated to each country's goods is computed as its 1983 GNP divided by the total GNP of the four countries in dollars. The shares are: U.S. 0.596, U.K. 0.079, Germany 0.111, and Japan 0.214. When residents of different countries are allowed to consume different baskets of goods, we use the consumption shares computed and reported in Frankel(1982, p. 272), which are derived from import and export to GNP ratios in 1973.

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