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Forward Speculation, Excess Returns, and Exchange Rate Variability: The Role of Risk Premiums

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Abstract

The paper reconsiders the unbiasedness hypothesis in the foreign exchange market. Within the context of a conventional model of exchange rates, risk premium shocks are constrained to have no permanent effects on the spot rate. Using monthly data from the post floating period, the paper estimates risk premiums for the Dollar rates of the Yen, Mark, and Pound. Risk premium innovations seem to explain a modest proportion of short term variability of exchange rate changes and excess returns. However risk premiums may explain serial correlations in excess returns.

1. Introduction

A common finding in international finance is that the forward rate is a biased predictor of the future spot rate. This bias is often explained by the existence of time-varying risk premiums and/or inefficient use of information by market participants. If agents are risk-neutral, then systematic discrepancies between exchange rate changes and the forward premiums are interpreted as evidence of irrationality on part of market participants. On the other hand, if market participants are rational but risk averse, the same systematic component can be attributed to a time-varying risk premium.

The controversy surrounding the relative size and variability of the foreign exchange risk premium took a new direction with Fama (1984). In investigating the role of risk premiums, Fama (1984) contends that the variance in the risk premium is reliably greater than the expected change

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in the exchange rate. Further work by Bilson (1985), Hodrick and Srivastava (1986), and Wolff (1987) confirm that a sizable portion of the bias is due to risk premiums. Frankel and Froot (1987), and Froot and Frankel (1989) use survey data for major currencies to decompose the source of bias into parts attributable to risk premiums and expectational errors. They find evidence of systematic expectational errors, and that forward premiums reflect expected depreciation rather than risk premiums. Recently, Chinn and Frankel (1994) extend the survey data set to include the newly industrializing countries and find evidence of biased expectations, which may be attributed to peso problems and learning behavior. In contrast to these studies that use survey data, Cavaglia et al. (1994) reexamines the bias in the forward premium and finds that the bias is due to both the failure of rational expectations and the existence of time-varying risk premium.

As a first step, it is important to examine the stochastic properties of excess returns in order to evaluate whether systematic expectational errors or risk premiums are responsible from the biased forecasts of the forward premium. Since a systematic excess return reflects unexploited profit opportunities, its time series behavior has an important bearing on market efficiency issues. Our departure from the previous literature is that we use a restriction consistent with a well-known model of exchange rate determination in identifying risk premiums. Specifically, the model implies that due to sticky prices and imperfect substitutability between domestic and foreign assets, the nominal exchange rate in the short run deviates from its long-run value. Using the strategy proposed by Blanchard and Quah (1989), one can identify the risk premium component of excess returns by restricting the risk premium shocks to have no long term effect on the spot rate.

We supplement conventional unit root tests such as Augmented Dickey-Fuller tests with the KPSS test developed by Kwiatkowski et al. (1992), which tests the null of stationarity against the alternative of a unit root. We also consider covariance stationarity against long memory in excess returns. To preview our results, we find that the excess return from forward speculation is covariance stationary. This result seems to be robust to various time series methods. By restricting risk premium shocks to have no long run effects on the spot rate, we identify risk premium sequences and obtain estimates of exchange rate variation due to risk premiums for the US Dollar rates of the British Pound, the Deutsche Mark (DM), and Japanese Yen. In all three cases, risk premium shocks account for less than 15% of the forecast error variance of exchange rate depreciation and excess returns.

Section 2 of the paper discusses issues related to risk premiums and presents a conventional model of exchange rate determination. The model is illustrative in that it has a simple formulation and it provides an identifying restriction that can be used within the structural Vector Autoregression (VAR) framework to recover risk premiums. Section 3 investigates time series properties of excess returns and presents our modeling strategy. Section 4 presents estimation results while Section 5 concludes.

2. Theoretical Framework

The existence of systematic discrepancies between the forward rate and the corresponding future spot rate led many to explain the discrepancy as evidence of a time varying risk premium.

Consider equation (1),

$$f_t = s_{t+1} + rp_t + \epsilon_{t+1} \tag{1}$$

where s_t is the spot rate of the foreign currency in units of domestic currency at time t , f_t is the forward exchange rate of the foreign currency quoted at time t for delivery at time $t+1$, rp_t is a risk premium, all variables are expressed in logarithms, and ϵ_{t+1} is an expectational error. It is evident from equation (1) that even if market participants are rational in the sense that the forecast error ϵ_{t+1} is orthogonal to the information set available at time t , risk aversion on part of market participants may lead to systematic departure of s_{t+1} from f_t .

A common specification for the hypothesis that the forward rate is an unbiased predictor of the future spot rate is the regression equation,

$$s_{t+1} - s_t = \alpha_0 + \beta_0(f_t - s_t) + u_{t+1} \quad (2)$$

where $(f_t - s_t)$ is the forward premium on foreign currency in units of domestic currency, and u_{t+1} is a random disturbance term. The restriction implied by absence of arbitrage is, $\alpha_0 = 0$, $\beta_0 = 1$, and lack of autocorrelation in u_{t+1} in case of non-overlapping forward contracts. Empirical evidence has overwhelmingly rejected the null hypothesis, and in most cases the estimate of β_0 is significantly less than zero. It is recognized that a biased β_0 can be due to a time-varying risk premium. An interesting approach to recover information regarding the risk premium is to consider the following regression specification due to Fama (1984):

$$f_t - s_{t+1} = \alpha_1 + \beta_1(f_t - s_t) + v_{t+1} \quad (3)$$

where regression equations (2) and (3) are complementary: $\alpha_0 = -\alpha_1$ and $\beta_0 + \beta_1 = 1$. Fama shows that the difference between the slope coefficients in equations (2) and (3) is

$$\beta_0 - \beta_1 = [\text{var}(rp_t) - \text{var}(s_{t+1}^e - s_t)] / \text{var}(f_t - s_t) \quad (4)$$

where an e over a variable denotes the expectational operator. Fama estimated this difference for nine major currencies and found it to be significantly positive. This can be interpreted as indicating a greater variance in the risk premium relative to the variance of the expected change in the exchange rate, which Fama contended was the main source of the estimated bias.

To understand the relationship between exchange rate changes, excess returns, and the forward premium consider the following decomposition of the change in the spot rate:

$$(s_{t+1} - s_t) = (s_{t+1} - f_t) + (f_t - s_t) \quad (5)$$

This equation highlights the conditions under which the forward premium will predict the change in the spot rate. If the excess return defined as $(s_{t+1} - f_t)$ in equation (5) is white noise, then the forward premium will be the best predictor of the depreciation rate, and we should expect $\beta_0 = 1$ in equation (2). With risk aversion, a covariance stationary excess return can be taken to represent the existence of a covariance stationary risk premium.¹ Note that the excess return is the sum of a risk premium and an expectational error as equation (1) implies. Since the innovations that affect expected depreciation affect excess returns as well, the forward premium will fail to predict the depreciation rate completely in this case. Moreover, since risk premiums are not observable, the statistical problem is to disentangle the observed excess returns into risk premiums and premium-free returns.

Risk Premiums and Exchange Rates: An Illustrative Model

In order to understand the relationship between the spot rate and risk premiums, consider the monetary/ portfolio balance synthesis model of exchange rate determination. The model extends the “overshooting” model of Dornbusch (1976) and Frankel (1979) to allow imperfect

substitutability between domestic and foreign bonds. In the model risk premiums and real interest differentials cause the exchange rate to deviate from its long-run equilibrium dictated by “macroeconomic fundamentals.” The long-run value of the exchange rate can simply be derived from long-run purchasing power parity (PPP), and a conventionally specified money market equilibrium. The long-run PPP states that:

$$\bar{s} = \bar{p} - \bar{p}^* \quad (6)$$

where p and p^* are the logarithms of domestic and foreign price levels, a bar over a variable signifies that the relationship holds in the long run, and for simplicity we omit time subscripts. The long run money market equilibrium is:

$$\bar{m} = \bar{p} + \phi\bar{y} - \lambda\bar{i} \quad (7)$$

where m is the logarithm of the money supply, y is the logarithm of real income, and i is the nominal interest rate. It is assumed that a similar relationship holds in the foreign money market.

Note that given covered interest arbitrage, an equivalent expression for the risk premium is:

$$rp = (i - i^*) - \Delta s^e \quad (8)$$

If the risk premium in the long run is zero, one can combine equations (6), (7), and (8) to obtain:

$$\bar{s} = (\bar{m} - \bar{m}^*) - \phi(\bar{y} - \bar{y}^*) + \lambda(\Delta\bar{p}^e - \Delta\bar{p}^{*e}) \quad (9)$$

where we eliminated the nominal interest differential in equation (8) using the fact that in the long run $\Delta s^e = \Delta p^e - \Delta p^{*e}$, and the risk premium is zero. With exogenous growth of output (or random

growth with mean zero), and a random walk monetary growth process, the rational expectations long-run equilibrium entails growth for the exchange rate and relative prices at the rate of current relative monetary growth $(\pi - \pi^*)$. In the short run, the exchange rate deviates from its long-run path and closes the gap with a speed of adjustment α :

$$\Delta s^e = \alpha(\bar{s} - s) + \pi - \pi^* \quad (10)$$

It is known that this specification for expectations is consistent with rational expectations; accordingly when the exchange rate lies on its equilibrium path, it is expected to increase at the rate, $(\pi - \pi^*)$. By adding and subtracting the nominal interest differential $(i - i^*)$, equation (10) can be rearranged to yield:

$$s - \bar{s} = -(1/\alpha) [(i - \pi) - (i^* - \pi^*)] + (1/\alpha)[i - i^* - \Delta s^e] \quad (11)$$

Equation (11) implies that the exchange rate will deviate from its long-run equilibrium value not only because of “sticky” commodity prices create a real interest differential as it is assumed in “overshooting” models, but also because a risk premium draws a wedge between domestic and foreign rates of return. Models along the portfolio balance framework specify the risk premium in equation (11) in terms of supply and demand for domestic and foreign assets using mean-variance optimization. Rather than deriving the reduced form for the risk premium and the exchange rate, we focus on the implied effects of the risk premium on the time series properties of the exchange rate.

It is evident from equations (9) and (11) that the risk premium has temporary effects on the spot rate. In the long run, the path of the exchange rate is determined by “macroeconomic

fundamentals.” This implication provides an important clue for our attempt to identify the risk premium sequences: restricting risk premium shocks to have no long run effect on the spot rate allows for the recovery of risk premiums from the observed excess return series. While the questionable performance of PPP may seem to invalidate this result, it is straightforward to incorporate deviations from PPP into the model as in Hooper and Morton (1982). When there are permanent deviations from PPP, our identification scheme is valid provided that deviations from PPP are not caused by risk premiums in the long run. Finally, there is an ample literature on the poor out-of-sample forecasting performance of conventional asset market models. We emphasize that our intention is not to test the model; rather, the theoretical model above serves to make the assumptions in our identification strategy explicit.

3. Time Series Properties of Excess Returns and Risk Premiums

Our data set consists of spot and 30-day forward rates of three commonly traded currencies: the Dollar rates of the DM, British Pound, and the Japanese Yen. The data are averages of bid-ask rates sampled at the close of trading on the last business day of the month from January 1974 through December 1995 taken from Data Resources Incorporated. Since the difference between two consecutive “last business days of the month” in general is not 30 days, a measurement error will be introduced due to data alignment. While data alignment and using transaction costs inherent in bid-ask spreads may seem to pose a problem, Bekaert and Hodrick (1993) construct data which incorporate costs inherent in bid-ask spreads and delivery structure of the market. They find essentially no difference in inference across specifications that are correctly constructed and those incorrectly specified.

Before attempting to identify risk premiums, it is important to examine the stochastic properties of the observed excess return $(s_{t+1}-f_t)$ series. First, we consider testing covariance stationarity of the returns using Augmented Dickey-Fuller (ADF) tests and KPSS tests due to Kwiatkowski et al. (1992), and the results are given in Table 1. The maximum lag length in the ADF tests is determined by starting with a maximum lag length and pairing down the lag length depending on the significance of the coefficient on the maximum lag. This method is proven to produce the true lag length with a unit asymptotic probability provided that the initial choice for lag length includes the true lag length. Table 1 indicates that the ADF test rejects the null hypothesis of a unit root for the three rates at the 5% significance level.

To check the robustness of the ADF test results, we use the KPSS test. This test takes stationarity as the null hypothesis, and is based on the LM score from the regression of the variable on a constant and possibly on a time trend. The statistic is given by

$$\hat{\eta}_\mu = \frac{\sum_{t=1}^{t=T} S_t^2}{T^2 s^2(l)} \quad (12)$$

where S_t^2 is the partial sum of the residuals from the regression, and $s^2(l)$ is a consistent estimate of the error variance corrected using the Bartlett window with the first l sample autocovariances, as in Newey and West (1987). The KPSS test was applied to each excess return series using lag truncations, $l = 0, 4, 8,$ and 12 . The results are given at the lower portion of Table 1.

The results of the KPSS test fail to reject the null hypothesis of stationarity for the three excess return series at conventional significance levels. Thus, KPSS tests confirm the evidence presented by ADF unit root tests.² Moreover, a time series plot of excess returns presented in Figure 1 does not indicate a trend in the mean or in the variance of excess returns.³

Table 1. ADF and KPSS Tests of Excess Returns

	Excess Returns ($s_{t+1}f_t$) in		
	Pound	DM	Yen
ADF Tests			
τ_μ statistic ^a	-4.46	-3.60	-3.25
KPSS Tests			
$\hat{\eta}_\mu$ statistic ^b			
lag truncation			
$l = 0$	0.130	0.284	0.115
$l = 4$	0.104	0.239	0.087
$l = 8$	0.102	0.232	0.083
$l = 12$	0.097	0.210	0.077

^a The critical values for the ADF tests at the 5 %, and 10 % significance levels are -2.88, and -2.58 respectively.

^b The upper tail critical values for KPSS tests at 5 % and 10 % levels are 0.463 and 0.347 respectively.

Identification of Risk Premiums: Methodology

Consider the excess return within the context of equation (1). Under rational expectations the *ex-ante* excess return for $(t, t+1)$, $er_t = (s_{t+1}f_t)$, can be written as the (negative) sum of a risk premium, rp_t and a white noise expectational error, ϵ_t . Given that excess returns are covariance stationary⁴, this can be interpreted as evidence of a covariance stationary risk premium. The model presented above implies that risk premiums have no long-run effect on the exchange rate. We use this identifying restriction to isolate the risk premium sequences.

Specifically, consider two types of orthogonal shocks that are the sources of variation in both Δs_t and er_t : a fundamental shock ϵ_{ft} and a shock to the risk premium ϵ_{rt} . The fundamental shock can be thought of as innovations due to “macroeconomic fundamentals” that affect the exchange rate in the long run. Using a strategy due to Blanchard and Quah (1989), the restriction that risk

premium shocks have temporary effects on the level of the spot rate is sufficient to enable the recovery of the whole risk premium sequence, rp_t . In addition, one can conduct innovation accounting exercises (impulse responses and variance decompositions) based on VARs to investigate the dynamic effects of the shocks.

More formally, we express the stationary stochastic processes Δs_t , and er_t as an infinite moving average process in the shocks ϵ_{ft} and ϵ_{rt} .

$$\Delta X_t = A_0 \epsilon_t + A_1 \epsilon_{t-1} + \dots = \sum_{i=0}^{\infty} A_i \epsilon_{t-i} \quad (13)$$

where $\Delta X_t = [\Delta s_t \ er_t]'$, $\epsilon_t = [\epsilon_{ft} \ \epsilon_{rt}]'$. An alternative notation is,

$$\begin{bmatrix} \Delta s_t \\ er_t \end{bmatrix} = \begin{bmatrix} a_{11}(L) & a_{12}(L) \\ a_{21}(L) & a_{22}(L) \end{bmatrix} \begin{bmatrix} \epsilon_{ft} \\ \epsilon_{rt} \end{bmatrix} = A(L) \epsilon_t \quad (14)$$

where $a_{ij}(L)$ are polynomials in the lag operator, L . The time paths of the effects of various shocks on the depreciation rate and the excess return are given by the coefficients of the polynomials $a_{ij}(L)$. Moreover, coefficient $a_{ij}^{(k)}$ in the $a_{ij}(L)$ polynomial is the response of variable i to a unit shock in ϵ_{jt} after k periods. The sum of all the moving average coefficients denoted $a_{ij}(1)$ gives the cumulative effect of ϵ_{jt} on variable i over time. The shocks can be normalized such that

$$E(\epsilon_t \epsilon_t') = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix} \quad (15)$$

In order to identify this model, one can estimate a finite order bivariate VAR:

$$\Delta X_t = B_1 \Delta X_{t-1} + \dots + B_k \Delta X_{t-k} + e_t \quad (16)$$

where the lag length is chosen such that residuals e_{it} , ($i = 1, 2$) approximate white noise, and

$$E(e_t e_t') = \Sigma = \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{bmatrix} \quad (17)$$

Since the elements of $\Delta \mathbf{X}_t$ are stationary, the system can be inverted to obtain the moving average representation:

$$\Delta X_t = e_t + C_1 e_{t-1} + C_2 e_{t-2} + \dots = \sum_{i=0}^{\infty} C_i e_{t-i} = C(L)e_t \quad (18)$$

The relationship between the orthogonal (pure) innovations ϵ_t and the composite innovations e_t is

$$e_t = A_0 \epsilon_t \quad (19)$$

Thus, the following relationship exists between the variance-covariance matrices:

$$E(e_t e_t') = A_0 E(\epsilon_t \epsilon_t') A_0' \quad (20)$$

and

$$\Sigma = A_0 A_0' \quad (21)$$

Since Σ is a symmetric matrix with known elements (or can be estimated consistently), it imposes 3 restrictions on the matrix \mathbf{A}_0 , which has 4 elements. An additional restriction is needed to identify \mathbf{A}_0 , so that the orthogonal shocks ϵ_{it} can be recovered using equation (19). Blanchard and Quah (1989) propose a long run neutrality restriction to achieve identification. If we evaluate the polynomials embedded in equations (13) and (18) at $L = I$, and note the relationship in equation (19), we obtain:

$$A(1) = C(1)A_0 \quad (22)$$

where $\mathbf{C}(1)$ contains known elements. The restriction that risk premium shocks have no long term effects on the spot rate corresponds to $a_{12}(1) = 0$ in equation (14) and provides the additional restriction needed to recover the elements of \mathbf{A}_0 . Once \mathbf{A}_0 is identified one can construct the \mathbf{A}_i

matrices of equation (13) as $\mathbf{A}_i = \mathbf{A}_0 \mathbf{C}_i$, $i = 1, 2, \dots$ and do variance decompositions and impulse response analysis typical of VARs based on the orthogonal shocks.

With the identifying restriction that $a_{12}(1) = 0$, we can estimate the bivariate system in equation (14) and construct the risk premium sequence as

$$rp_t = a_{22}(L)\epsilon_{rt} = \sum_{k=0}^t a_{22}^{(k)}\epsilon_{rt-k} \quad (23)$$

Furthermore, estimation of the system in (14) allows us to assess the relative contributions of “risk premium” and “fundamental” shocks to the depreciation rate and excess returns.

4. Empirical Results

We follow the identification strategy outlined above and estimate the finite order VAR in equation (16) with 6 lags for the Pound and DM rates, and with 10 lags for the Yen rate.

Likelihood ratio tests and residual diagnostics indicate that these lag lengths are sufficient to capture the dynamics embedded in the data. After estimating the system as an infinite moving average process in the orthogonal shocks, we present variance decomposition results in Table 2.

The results indicate that risk premium shocks account for between 12 % and 14 % of the forecast error variance of the depreciation rate in the Pound rate, up to 2.5 % in the DM rate and between 9 % and 11 % of the variance in the Yen depreciation rate. The rest of exchange rate variability can be thought of as due to the arrival of “news” about “fundamentals.” As for excess returns, risk premium shocks account for between 10-13 % of the variance in excess returns in the Pound rate, 7-10 % of the variance in the Yen rate, and less than 2.5 % in the DM rate. Note that after the first month the proportion of the forecast error accounted for by risk premiums does not

Table 2. Variance Decompositions

attributable to	Proportion of the Forecast Error in				
	ϵ_f	Δs_t	ϵ_r	er_t	
<i>Pound</i>					
Horizon					
1-month	87.7	12.3	90.2	9.8	
3-month	86.9	13.1	88.3	11.8	
12-month	86.1	13.9	87.1	12.9	
24-month	86.1	13.9	87.0	13.0	
<i>DM</i>					
Horizon					
1-month	98.7	1.3	99.2	0.8	
3-month	98.2	1.8	99.0	1.0	
12-month	97.6	2.4	97.9	2.1	
24-month	97.5	2.5	97.6	2.4	
<i>Yen</i>					
Horizon					
1-month	90.5	9.5	92.5	7.5	
3-month	90.2	9.7	92.3	7.7	
12-month	89.3	10.6	90.9	9.1	
24-month	89.0	11.0	90.3	9.7	
<i>Long Run Impact Multipliers</i>					
		$a_{11}(1)$	$a_{12}(1)$	$a_{21}(1)$	$a_{22}(1)$
Pound		0.0342	0	0.0367	0.0107
DM		0.0326	0	0.0351	0.0142
Yen		0.0361	0	0.0371	0.0074

increase significantly. Overall, risk premiums account for a modest proportion of exchange rate variability and the observed discrepancies between forward rates and the corresponding spot rates. Our results indicate a smaller role of risk premiums than those reported by Fama (1984) and Wolff (1987).

The estimates of the long-run impact multipliers are given at the lower part of Table 2. The multipliers give the *cumulative* response of each variable (the depreciation rate and the excess return) to respective shocks. For example, a typical unit shock to the fundamental raises the

Pound spot rate by 3.42 % , the DM rate by .26 % , and the Yen rate by 3.61 % in the long run. Note that the long-run impact of the fundamental shock on excess returns exceeds the impact on the depreciation rate. This can explain the observed empirical regularity that the observed variability of the excess return is greater than the depreciation rate. This can be seen by noting that the variance of each orthogonal shock is normalized to unity and risk premium shocks have no cumulative effect on the depreciation rate, i.e., $a_{12}(1) = 0$, so that

$$\begin{aligned} \text{var}(\Delta s_t) &= a_{11}(1)^2 \\ \text{var}(er_t) &= a_{21}(1)^2 + a_{22}(1)^2 \end{aligned} \tag{24}$$

where the relations follow from equation (14) above. Given that the fundamental shock has similar long-run impacts on both the excess return and the depreciation rate (see the long-run impact multipliers in Table 2), the risk premium acts to magnify the variability of the return. In each case, the estimated variability of excess returns exceeds that of the depreciation rate. Our results indicate that the observed high variability of excess returns relative to the depreciation rate may be consistent with a low variability of risk premiums relative to the depreciation rate.

Statistical Properties of Risk Premiums and Risk-Premium-Free Excess Returns

Using equation (23), we construct the risk premium sequence for each exchange rate and the time series plots of risk premiums along with excess returns are given in Figure 1. It is apparent from the figure that risk premiums account for a smaller proportion of excess returns. The figure also conforms to variance decomposition results in that the variability of risk premiums in the Yen and Pound rates is higher than that of the DM rate.

Having estimated risk premium sequences, we construct the risk premium free excess returns

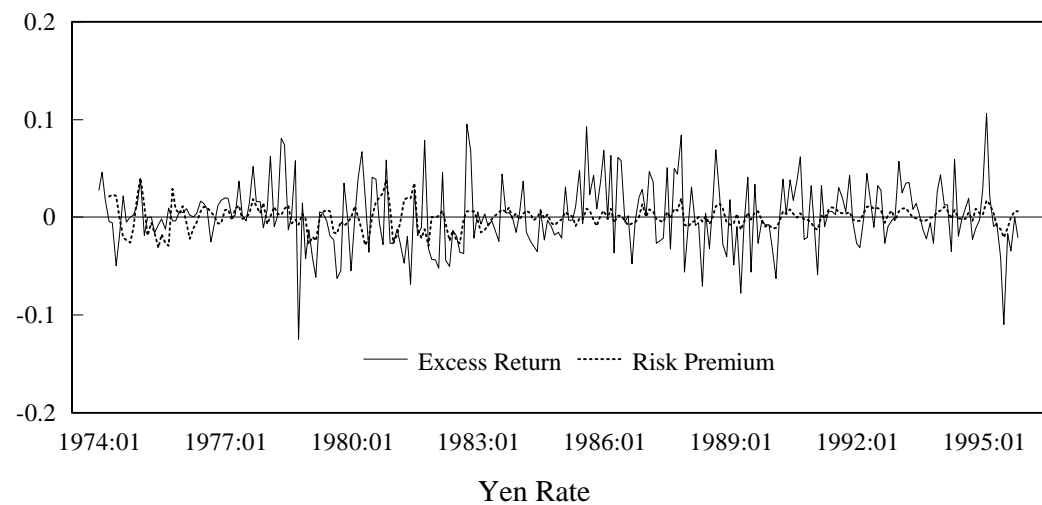
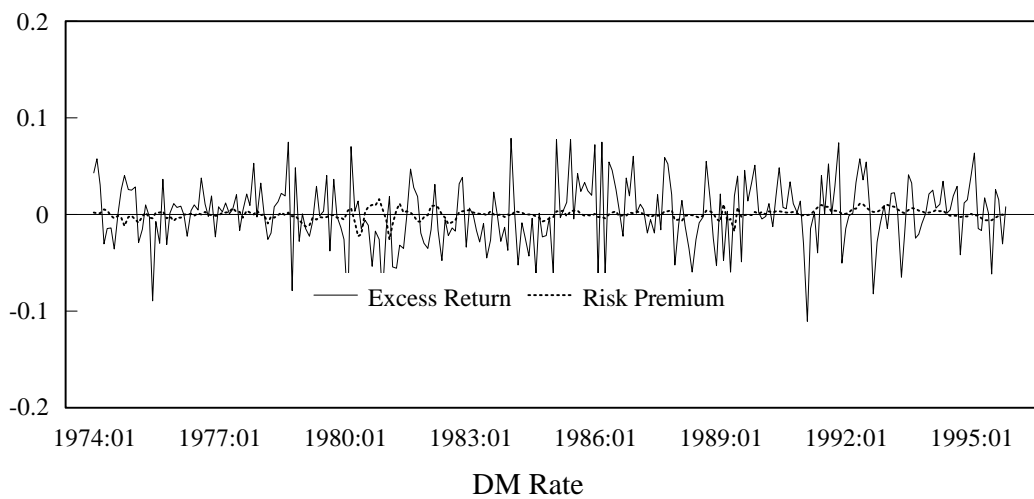
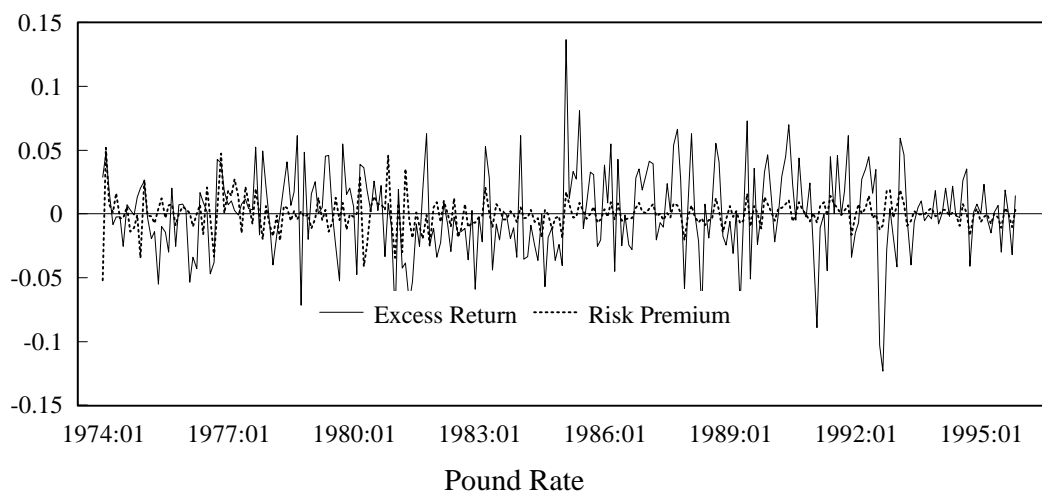


Figure 1. Excess Returns and Risk Premia

as $(s_{t+1}-f_t)-rp_t$ and present descriptive statistics for risk premiums and premium-free excess returns in Table 3. The mean risk premium is zero for the three rates. We note that theoretical models of international asset pricing do not impose sign restrictions on risk premiums (Adler and Dumas, 1983). Moreover, the risk premium sequences are small in size and variability relative to the observed excess return series. This is in line with previous studies which derived risk premiums from theoretical asset pricing/portfolio choice models. For example, Macklem (1991) simulated a version of the Lucas (1982) model, which is a consumption based representative agent model of asset pricing, and found that the produced risk premiums are too small. Similarly, Bekaert (1994) used a cash-in-advance model and found the risk premium to be too small to account for the unbiasedness hypothesis.

Table 3 indicates that the measures of risk premium sequences in the Pound and Yen rates have roughly twice the variability in the DM rate. Variance decomposition results also showed a larger role of risk premiums in the Yen and Pound rates relative to the DM rate. A visual inspection of Figure 1 indicates that the risk premium for the Yen and Pound rates is more variable prior to 1982 while the premium for the DM rate seems to be relatively smooth for the entire estimation period. It is interesting to note that Germany has removed most barriers to capital mobility in 1974 (Dooley and Isard, 1980), while the UK and Japan liberalized their financial markets in late 1970s and early 1980s. For example Kasman and Pigott (1988) show that between 1974-79, barriers to capital mobility as reflected in onshore-Euromarket interest differential were considerably higher for the UK and Japan than for Germany.

In order to get additional insight and consider the possible effects of structural change, we re-estimated the model for 1981-95. The results are surprising in that risk premiums account for a

Table 3. Statistical Properties of Risk Premia and Premium-Free Excess Returns

	Pound	DM	Yen
<i>Risk premium</i>			
Mean	0.000	0.000	0.000
s.d.	0.012	0.005	0.011
J-B ^a	192.94***	223.73***	14.53***
<i>Cross-correlations of risk premia</i>			
Yen	0.26	0.49	
DM	0.31		
<i>Premium-free return: $(s_{t+1}f_t)-rp_t$</i>			
Mean	0.001	0.001	0.001
s.d.	0.032	0.034	0.032
J-B	12.02***	5.68*	6.61**
Q(2) ^b	7.97**	3.26	3.17
Q(4)	8.42*	3.59	4.71
Q(8)	10.93	7.53	10.64
Q(16)	19.76	15.07	19.78
Q(30)	33.19	27.64	34.33
Q(40)	44.25	36.55	41.25
<i>Premium-free return: $(s_{t+1}f_t)-rp_{t-1}$</i>			
Q(2)	0.09	2.69	0.15
Q(4)	1.33	2.78	1.52
Q(8)	2.66	6.73	5.62
Q(16)	15.67	14.69	15.83
Q(30)	24.18	25.24	32.20
Q(40)	31.96	33.98	41.85

^a J-B is the Jarque-Bera statistic for testing normality. Rejection of the null hypothesis is indicated by (***) at the 1 %, (***) at the 5 %, and (*) at the 10 % level.

^b Q(k) gives the Ljung-Box statistic for up to kth order serial correlation. Rejection is indicated by (**) at the 5 %, and (*) at the 10 % level.

higher proportion of excess returns in the Yen and Pound rates while they still account for a

negligible proportion (less than 2%) in the DM rate.⁵ A possible explanation can be found in the

conduct of macroeconomic policies in these countries. Using data from September 1982 - October 1996, for example, Frankel and MacArthur (1988) show the *ex ante* and *ex post variability* in the local vs. Eurodollar real interest rate differential to be higher for Japan and the UK than for Germany. While this evidence in itself may not provide a satisfactory explanation, it may suggest the existence of some common factors that can explain both the differential incidence of risk premiums and real interest rate variability in the three countries.

A possibility for the low incidence of risk premiums in the DM rate may be credibility issues surrounding the European Monetary System (EMS).⁶ A substantial literature articulated within the context of the EMS alludes to fixed exchange rates as providing discipline to policymakers; accordingly fixed rates can be viewed as a commitment mechanism that prevents the policymaker from resorting to inflationary policies (Giavazzi and Pagano, 1988). As such, EMS commitments may have restrained Germany from pursuing a higher degree of discretionary policies. Clearly, the relationship between foreign exchange risk premiums and the conduct of macroeconomic policy is an important question which goes beyond the scope of the current paper.

Note that for each rate, the Jarque-Bera (J-B) test indicates a strong rejection of normality for risk premiums. The estimates of cross-currency correlations of premiums are in the range 0.26-0.49 indicating that risk premium sequences for different currencies relative to the US Dollar have not moved much together. Risk premium free excess returns have a mean of 0.1 % and exhibit much higher variability than risk premiums. Except for the Pound rate, there is no strong evidence of serial correlation in premium-free excess returns. If one interprets the premium-free excess return as due to expectational errors, it is not clear whether this represents *ex ante* unexploited profit opportunities in the Pound rate. Moreover, it is interesting to note that if one constructs the

premium-free returns using the lagged risk premium, the serial correlation completely disappears. This seems to be robust across the three currencies considered. This may possibly reflect a learning behavior on part of market participants. Since excess returns are not observable at time t , agents may use information on the lagged premium when they take positions in the market.

Finally, autocorrelations in risk premiums and premium-free excess returns are given in Figure 2. As in Wolff (1987) the estimates of premiums show a high degree of persistence, and seem to be behind the serial correlations in excess returns. Although premium free excess returns show a few significant serial correlations, these again disappear if one considers the lagged premium-free returns. Overall the results seem to be broadly consistent with the notion of an efficient market.

5. Conclusions

We have investigated stochastic properties of excess returns from forward speculation for Dollar rates of the Pound, Mark, and Yen for the period January 1974 through December 1995 and found evidence that it is covariance stationary. This result seems to be robust to various time series methods. This may be interpreted as evidence of risk premiums in the short run. In order to identify risk premiums, we restrict risk premium shocks to have no long-run effect on the level of the spot rate. We also show that this restriction is consistent with a popular model of exchange rate determination. We then investigate the properties of the identified risk premiums and premium-free returns.

Variance decomposition results indicate that risk premium shocks account for a modest proportion of spot rate and excess return variability. Risk premiums seem to be strongly serially correlated, and account for the serial correlation in excess returns reasonably well. When excess

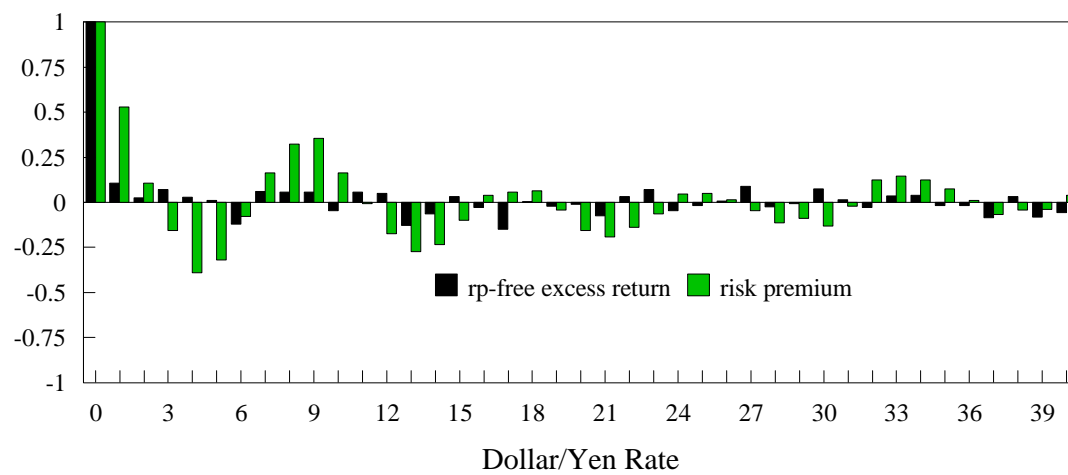
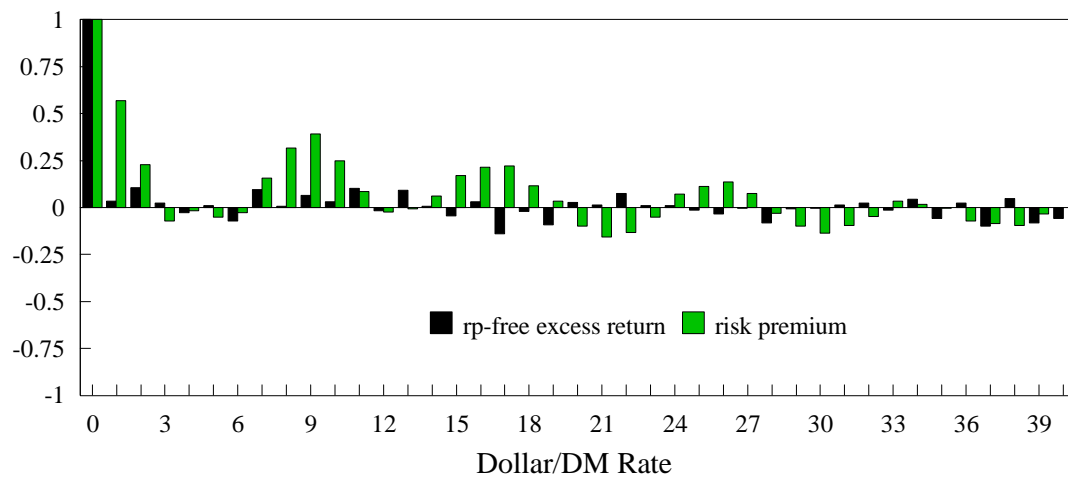
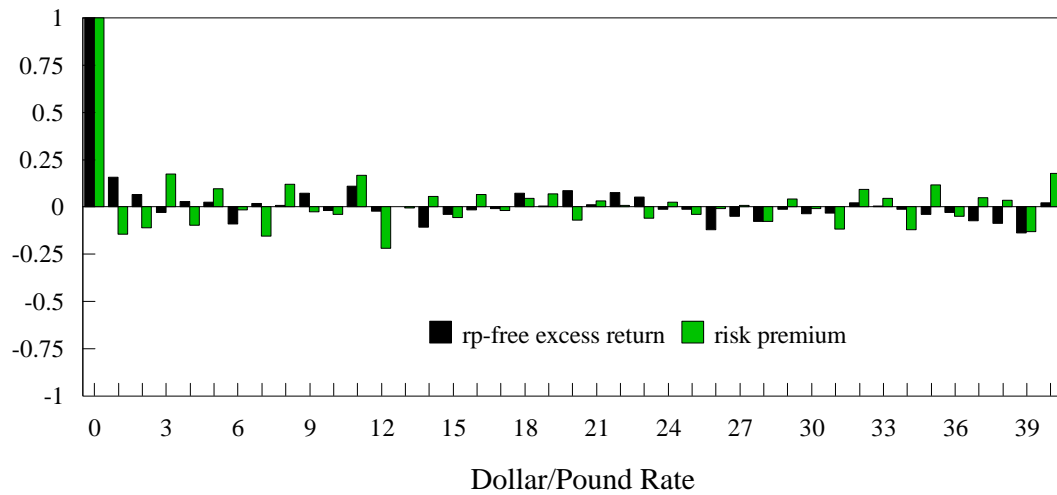


Figure 2. Autocorrelations in Risk Premia and Risk Premium Free Excess Returns

returns are filtered with risk premiums, the serial correlation in excess returns is reduced considerably. Interestingly, when excess returns are filtered using the lagged risk premium with a single lag, the serial correlation in excess returns disappears. Overall, our results are broadly consistent with market efficiency.

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Endnotes

1. If the forward bias is not covariance stationary, then there is a permanent component in the sequence of returns from forward speculation. In this case, it may not be meaningful to assume the existence of risk premiums.
2. Meese and Singleton (1982) used ADF tests to examine forward market efficiency and concluded that the excess return is a unit root process. However, Corbae and Ouliaris (1986) accounted for overlapping contracts and rejected the unit root hypothesis in excess returns.
3. In addition to the unit root tests, we examined the low frequency behavior of excess returns by estimating the power spectrum for the excess return series. The spectrums showed no trend at low frequencies. Although KPSS tests are quite powerful against fractionally integrated alternatives (Lee and Schmidt 1993), we also estimated the fractional differencing parameter d , using a semi-parametric method due to Geweke and Porter-Hudak (1983). The estimates of d range from $d = 0.02$ for the Yen and $d = 0.04$ for the Pound, to $d = 0.09$ for the DM. These estimates are based on 16 low-frequency ordinates (square-root of the number of observations); however the results do not seem to be sensitive to the number of low frequency ordinates. The results confirm covariance stationarity and are not reported for brevity.
4. Covariance stationarity by itself is not suggestive of risk premiums; we examine the autocorrelation functions of excess returns, and find that there is evidence of significant serial correlation particularly in the Yen and Pound rates.
5. Results are not reported for space considerations and are available upon request.
6. The British participation in the Exchange Rate Mechanism of the EMS was brief: between October 1990 and September 1992, and at a larger band ($\pm 6\%$) than the usual ($\pm 2.5\%$).