RETHINKING DEVIATIONS FROM UNCOVERED INTEREST PARITY: THE ROLE OF COVARIANCE RISK AND NOISE*

Nelson C. Mark and Yangru Wu

We examine the ability of the standard intertemporal asset pricing model and a model of noise trading to explain why the forward foreign exchange premium predicts the future currency depreciation with the 'wrong' sign. We find that the intertemporal asset pricing model is unable to predict risk premia with the correct sign to be consistent with the data. The noise-trader model, while highly stylised, receives fragmentary support from empirical research on survey expectations.

This paper investigates the theoretical basis of an asset pricing anomaly in international finance known as the forward premium bias. That is, the empirical finding that the forward premium helps to predict the future percentage rate of currency depreciation but not with the sign implied by uncovered interest parity.[1] As a corollary to the forward premium bias, Fama (1984) demonstrates that the deviation from uncovered interest parity, which we denote by p_t, is negatively correlated with and is more volatile than the rationally expected rate of depreciation. These empirical findings have long posed a challenge to international economic theory. In this paper, we explore two very different approaches to explaining this puzzle: the standard representative-agent intertemporal asset pricing model and a model of noise trading.

Our analysis of the intertemporal asset pricing model employs quarterly data for the United States, Great Britain, Germany, and Japan extending from 1976.1 to 1994.1 and asks if p_t can plausibly be viewed as a risk premium. The model is evaluated by examining how well the risk premia implied by the theory match up with statistical estimates of p_t. Preferences of the representative agent are assumed to possess a simple form of habit persistence while our statistical estimates of p_t are derived from a VECM (vector error correction model) fitted to the logarithms of the spot and forward exchange rates. The main finding of this section is that the model fails to predict p_t with the correct sign. The recent literature has emphasised the model's ability to match the volatility in forward foreign exchange rate returns but our results suggest that

[1] See Engel (1996), Froot and Thaler (1990), Hodrick (1987), and Lewis (1995) for surveys of this literature.

the failure of the model in understanding the data occurs at a more primitive level.[2]

Our second exploration examines the De Long *et al.* (1990) model which combines fundamentalist investors whose expectations are rational, with noise traders whose beliefs concerning future investment returns are distorted. In this model, the heterogeneity in beliefs among economic agents creates the basis for trading volume and induces systematic movements in p_t that are correlated with the forward premium. The p_t that emerges in this model is unrelated to covariance risk in the usual sense. Instead, short-horizon fundamentalist traders bear the risk in that they may be required to liquidate their positions at a time when noise-traders have pushed asset prices even farther from their fundamental values than they were when the investments were formed.

Our specification of noise-trader beliefs is guided in part by the evidence, reported by Froot and Frankel (1989) in their econometric study of survey expectations, that professional foreign exchange market participants place excessive weight on the forward premium when forming expectations of the future depreciation. In addition to providing a potential explanation for the forward premium bias, the noise–trader model supplies an account of the apparent short-term overreaction of exchange rate changes and the gradual adjustment towards its fundamental value in the long run that has been reported in other empirical research.[3] Finally, we argue that empirical estimates drawn from the extant literature on survey expectations of foreign exchange forecasts provides fragmentary support for the model.

We note that the call for an international transactions tax on foreign exchange to inhibit short-term capital flows (Eichengreen *et al.*, 1995) is an important policy issue on which our research has some bearing. The underlying premise for the tax is that frequent swings in investor sentiment occurring independently of changes in economic fundamentals generate financial market volatility with undesirable externalities. The noise-trader model that we examine represents such an environment. To sound a note of caution, however, we grant that many might find our model to be too stylised to legislate on whether such a transactions tax should be implemented. Naturally, a proper assessment of this issue would require a careful evaluation of the entire body of research on foreign exchange market efficiency, of which the present paper is small part.

The paper is organised as follows. The next section begins by presenting our estimates of the forward premium bias and of p_t. Section 2 asks if statistical estimates of p_t behave like the risk premia of the intertemporal asset pricing

---

[2] See Backus *et al.* (1993), Bekaert (1994) and Cecchetti *et al.* (1994) who study the model along the volatility dimension.

[3] Mark (1995), Chinn and Meese (1995), Mark and Choi (1997), Chen and Mark (1996), and Lothian and Taylor (1996) report empirical evidence of long-horizon reversion of exchange rates to their fundamental values. In related work on quasi-rational modelling of exchange rate determination, Goldberg and Frydman (1996) show that the exchange rate will overshoot and drift away from the fundamentals when agents hold heterogeneous beliefs and have imperfect knowledge of the economy.

model. The noise trader model is presented in section 3, and section 4 concludes.

1. The Forward Premium Bias

The exchange rate is the domestic currency price of one unit of the foreign currency. Throughout the paper, upper case letters will denote levels of variables and with the exception of interest rates, lower case letters will denote their respective logarithms. In the empirical work, logarithms are multiplied by 100 so as to state differences in percentage terms. Now at date t, let s_t be the spot exchange rate and f_t be the forward exchange rate for date $t+1$ delivery. Let $x_t \equiv f_t - s_t$ be the forward premium, which is equal to the nominal interest differential between domestic and foreign currency debt by virtue of covered interest parity. To understand the behaviour of p_t we begin with Fama's (1984) decomposition,

$$x_t = \mathrm{E}_t(\Delta s_{t+1}) + p_t. \tag{1}$$

From (1), the *ex post* depreciation can be expressed as $\Delta s_{t+1} = x_t + (v_{t+1} - p_t)$, where $v_{t+1} = \Delta s_{t+1} - \mathrm{E}_t(\Delta s_{t+1})$ is the rational forecast error. It is the presence of p_t that generates the forward premium bias. Since p_t will, in general, be correlated with x_t, the regression,

$$\Delta s_{t+1} = \alpha + \beta x_t + \epsilon_{t+1}, \tag{2}$$

will be subject to an omitted variables bias resulting in estimates of the slope coefficient that deviate from 1.[4] Our own estimates of (2) from a quarterly sample of USD (dollar) rates of the GBP (pound), the DEM (deutsche-mark) and the JAY (yen), as well as three cross rates are reported in Table 1.[5]

As can be seen in the full sample estimates, the slope coefficients are all negative. Under one-sided tests, they are significantly so at the 5% level for the USD/GBP rate and at the 1% level for the USD/JAY and GBP/JAY. The full-sample SUR (seemingly unrelated regression) estimate of the slope is -1.39 ($s.e. = 0.37$), which is significantly negative at the 1% level.

Following Bekaert and Hodrick (1993), we break the sample at 1980.I and confirm their finding that the forward premium bias seems to be dominated by the 1980s. The analysis of the first sub-sample which extends from 1976.I–1979.IV yields positive OLS slope-coefficient estimates for the USD/GBP and GBP/DEM rates. The remaining OLS estimates and the SUR estimate of β, while negative, are not significant. Estimates from the second subsample,

[4] Because s_t and f_t appear to be I(1), researchers such as Evans and Lewis (1993) argue that the forward bias should be measured by the cointegrating regression $s_{t+1} = \alpha + \beta f_t + v_{t+1}$ since under the alternative ($\beta \neq 1$), $\mathrm{E}_t s_{t+1} = \beta f_t$ and the error term in (2) is $\epsilon_{t+1} = (\beta - 1) s_t + v_{t+1} \sim$ I(1). We study (2) because we are interested in studying the behaviour of a stationary p_t and how its presence biases the slope coefficient away from 1. We take as a maintained hypothesis that the forward premium, and therefore the expected excess return p_t are I(0).

[5] The sample extends from 1976.1 to 1994.1. We follow Hansen and Hodrick (1983) by starting the sample in 1976.1 after the Rambouillet Conference. The sources for the exchange rates as well as the other data used in the paper are described in the Appendix.

Table 1

Regressions of Quarterly Depreciation on 3-Month Forward Premium

$\Delta s_{t+1} = \alpha + \beta x_t + \epsilon_{t+1}$

	USD/GBP	USD/DEM	USD/JAY	GBP/DEM	GBP/JAY	DEM/JAY	$\hat{\beta}_{SUR}$
			(a) 1976:I–1994:I				
$\hat{\alpha}_{OLS}$	−1.340	0.638	3.294	1.622	7.702	1.041	
	(0.895)	(0.886)	(0.964)	(1.116)	(1.687)	(0.648)	
$\hat{\beta}_{OLS}$	−1.522	−0.136	−2.526	−0.602	−4.261	−0.755	
	(0.863)	(0.839)	(0.903)	(0.782)	(1.133)	(1.042)	
$\hat{\alpha}_{SUR}$	−1.252	1.324	2.475	2.577	3.728	1.151	−1.390
	(0.721)	(0.784)	(0.757)	(0.736)	(0.820)	(0.618)	(0.368)
			(b) 1976:I–1979:IV				
$\hat{\alpha}_{OLS}$	2.346	2.614	2.948	0.546	6.366	−1.248	
	(1.511)	(1.946)	(1.853)	(3.509)	(4.326)	(1.693)	
$\hat{\beta}_{OLS}$	1.247	−0.034	−1.967	0.464	−3.621	−0.897	
	(1.162)	(1.848)	(1.563)	(1.802)	(2.362)	(3.956)	
$\hat{\alpha}_{SUR}$	0.912	2.851	1.688	1.938	0.775	−1.163	−0.298
	(1.270)	(1.140)	(1.456)	(1.716)	(2.044)	(1.490)	(0.693)
			(c) 1980:I–1994:I				
$\hat{\alpha}_{OLS}$	−2.132	0.318	3.433	1.934	8.148	1.680	
	(1.040)	(0.997)	(1.148)	(1.218)	(2.021)	(0.716)	
$\hat{\beta}_{OLS}$	−2.419	−0.199	−2.669	−0.980	−4.549	−1.311	
	(1.079)	(0.951)	(1.119)	(0.982)	(1.471)	(1.070)	
$\hat{\alpha}_{SUR}$	−1.694	0.968	2.746	2.662	4.440	1.777	−1.684
	(0.847)	(0.927)	(0.886)	(0.802)	(0.891)	(0.658)	(0.438)

Notes: Standard errors in parentheses. Units are in percent. Δs_t and x_t are changes in the logarithms of the spot rate and the 3-month forward premium rate multiplied by 100.

extending from 1980.I–1994.I, yields OLS slope-coefficient estimates that are all negative and significantly so for three of the exchange rates (USD/GBP, USD/JAY and GBP/JAY). The SUR estimate is $\hat{\beta} = -1.684$ (s.e. = 0.438) over this period. The results of the table are at odds with uncovered interest parity, which implies that the forward premium is an unbiased predictor of the future depreciation, $(\beta = 1)$. While the forward premium is found to help in prediction, it does so with the wrong sign.[6]

Fama (1984) deduced additional properties of p_t. First, he showed that the negative slope coefficients from these regressions imply that p_t is negatively correlated with $E_t \Delta s_{t+1}$. Second, he demonstrated that p_t is more volatile than $E_t \Delta s_{t+1}$, suggesting that it is potentially quite large. One theory about p_t is that it is a risk premium.[7] We now examine the ability of the intertemporal asset pricing model to generate plausible risk premia in accordance with the data.

[6] In a recent study, Wu and Zhang (1997) report that the forward premium bias is robust to two distribution-free non-parametric tests.

[7] It is unlikely that p_t is the result of a no-trading band arising from transactions costs. Bekaert and Hodrick (1993) and Breuer and Wohar (1996) carefully sample the data by accounting for the 2-day delivery lag on spot and forward transactions and by using appropriate bid and ask prices in calculating returns. They find that making these adjustments has little importance in explaining the forward premium bias.

2. Does p_t Compensate for Covariance Risk?

The analysis of this section compares time-series estimates of p_t to risk premia generated by an economic model of risk. We begin this section by describing the statistical estimates of p_t that we use.

2.1. *Time-Series Estimates of p_t*

Our estimates of p_t were obtained by modelling and estimating s_t and f_t as a bivariate VECM where the forward premium was constrained to be stationary; i.e., the cointegration vector was set at $(-1, 1)$ *a priori*.[8]

The VECM estimates of p_t for USD rates of the GBP, DEM, and JAY, with their 2-standard-error bands are presented in Figs. 1a–1c. The Figs. also display our point estimates of $E_t(\Delta s_{t+1})$. The estimated series are seen to be persistent, especially for the GBP and JAY. Both $\hat{p}_t$ and $\hat{E}_t(\Delta s_{t+1})$ alternate between positive and negative values and change sign infrequently. Significantly positive and negative values are found for $\hat{p}_t$. We note further that correlated movements across the currencies are visable: Each of the series contain spikes in early 1980 and 1981, the $\hat{p}_t$s are generally positive during the period of dollar strength from mid-1980 to 1985 and are generally negative from 1990 to late 1993. Also evident is the negative covariance between $\hat{p}_t$ and $\hat{E}_t(\Delta s_{t+1})$.

The next section compares the behaviour of our time-series estimates of p_t to risk premia generated by the intertemporal asset pricing model.[9]

2.2. *Implications from Euler Equations*

At date t, let R_t be the gross nominal return on a domestic currency denominated one-period discount bond, F_t be the one-period forward exchange rate, S_t be the spot exchange rate, C_t be consumption expenditures of the representative agent, and π_t be the purchasing power of the domestic money (the reciprocal of the price level). We study a class of utility functions examined by Hansen and Jagannathan (1991), in which period utility displays CRRA (constant relative risk aversion) defined over consumption services, $C_t^s = C_t + \delta C_{t-1}$ produced at date t. Expected utility is the expected discounted sum of period utilities,

$$U_t = E_t \sum_{k=0}^{\infty} \theta^k \frac{(C_{t+k}^s)^{1-\gamma} - 1}{(1 - \gamma)},\tag{3}$$

where γ is the coefficient of relative risk aversion and θ is the subjective discount factor. Equation (3) is the standard CRRA formulation when $\delta = 0$. Consumption purchases contain a durable component when $\delta > 0$, and

---

[8] The details of estimation and diagnostic evaluation of the VECM are suppressed in the interests of conserving space but are available upon request. We simply note here that the VECM that we estimated provides a reasonably accurate description of the salient properties of the data.

[9] Much research has been devoted to estimating p_ts. See Cumby (1988), who employs a projection procedure, Cheung (1993), Wolff (1987), and Hai *et al.* (1997) who utilise Kalman filtering techniques, and Bekaert and Hodrick (1992) and Canova and Ito (1991) who exploit VAR methods.
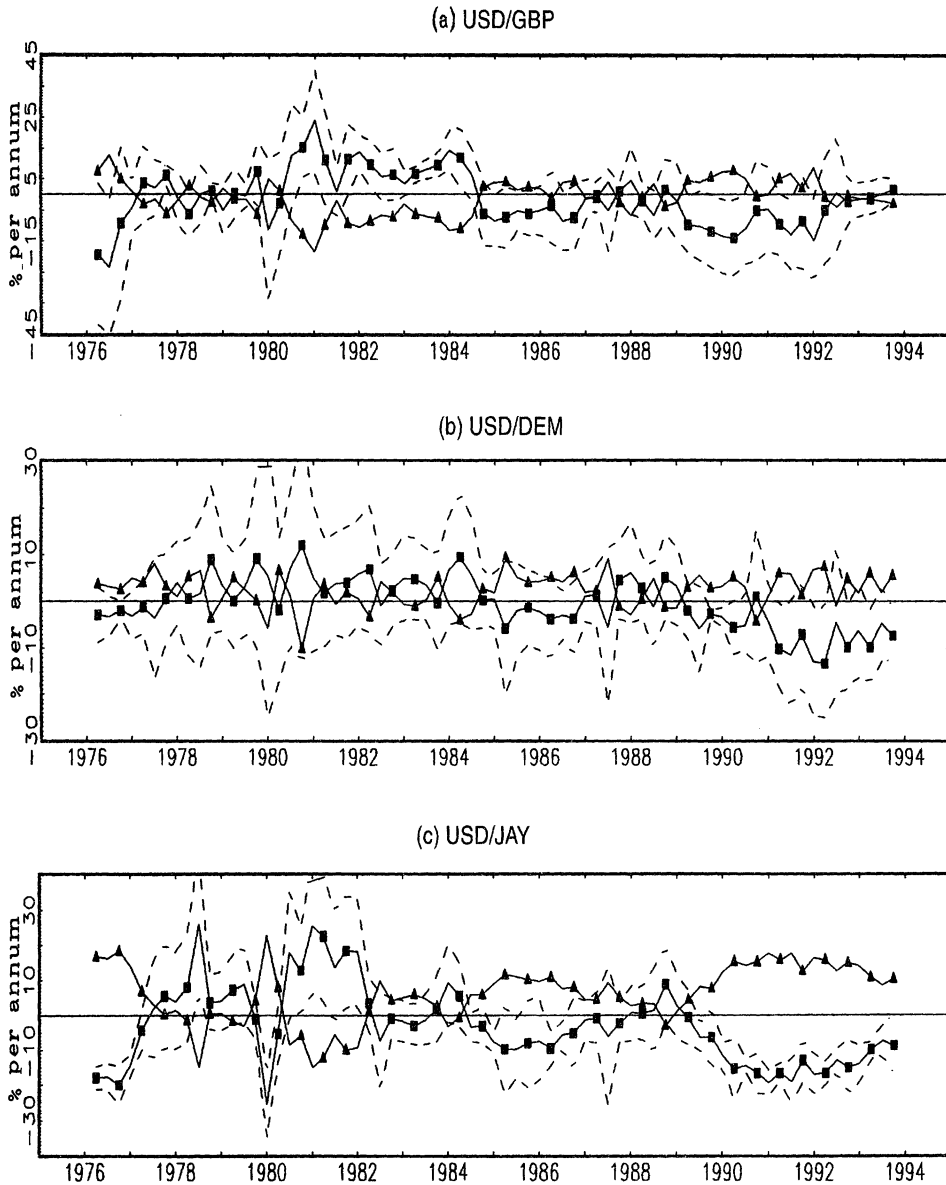
(a) USD/GBP

(b) USD/DEM

(c) USD/JAY

Fig. 1. *Time series point estimates of p_t (boxes) with 2-standard error bands and point estimates of $E_t(\Delta s_{t+1})$ (triangles). (a) USD/GBP, (b) USD/DEM, (c) USD/JAY.*

preferences exhibit a form of habit persistence when $\delta < 0$. The intertemporal marginal rate of substitution of money for these preferences is,

$$m_{t+1}(\theta, \gamma, \delta) \equiv \theta \frac{(C^s_{t+1})^{-\gamma} + \delta\theta(C^s_{t+2})^{-\gamma}}{(C^s_t)^{-\gamma} + \delta\theta(C^s_{t+1})^{-\gamma}} \frac{\pi_{t+1}}{\pi_t}.$$

To economise on notation, let $\tilde{s}_{t+1} \equiv (F_t - S_{t+1})/S_t$ denote the speculative profit on forward foreign exchange. The Euler equations for pricing the forward exchange risk premium and a domestic currency bond are, respectively,

$$0 = E_t[m_{t+1}(\theta, \gamma, \delta)\tilde{s}_{t+1}], \tag{4}$$

$$\frac{1}{R_t} = E_t[m_{t+1}(\theta, \gamma, \delta)]. \tag{5}$$

We note that $E_t(\tilde{s}_{t+1})$ differs from p_t by a Jensen's inequality term. Since the weight of the available evidence suggests that it is empirically unimportant, we ignore the effects of Jensen's inequality and combine (4) and (5) to obtain,[10]

$$p_t = E_t(\tilde{s}_{t+1}) = -R_t \text{Cov}_t[m_{t+1}(\theta, \gamma, \delta), \tilde{s}_{t+1}]. \tag{6}$$

According to (6), the home currency is risky when $\text{Cov}_t[m_{t+1}(\theta, \gamma, \delta), \tilde{s}_{t+1}]$ is negative. That is because the value of the home currency tends to be low at the same time that consumption is low. The home currency thus serves as a poor hedge against bad states of nature. One of the sharpest implications of the model is that p_t and $\text{Cov}_t[m_{t+1}(\theta, \gamma, \delta), \tilde{s}_{t+1}]$ have opposite signs. Noting that p_t is in the date t information set, $m_{t+1}(\theta, \gamma, \delta)$ and $\tilde{s}_{t+1}$ should be negatively correlated whenever preceded by positive values of p_t.

We test these sign restrictions by sorting paired observations of $(m_{t+1}(\theta, \gamma, \delta), \tilde{s}_{t+1})$ into two groups according to whether they were preceded by positive or negative values of our estimates of p_t. The sorting was done from the perspective of a US individual who speculates against the GBP, DEM, and JAY and from the perspective of UK, German, and Japanese investors who speculate against the USD. In making these calculations, we need to specify values for (θ, γ, δ) in advance. To put the theory in the best possible light, we report results under habit persistence with $\theta = 0.99$, $\gamma = 7$, $\delta = -0.5$. Analogous calculations performed under CRRA utility ($\delta = 0$) and under durability ($\delta = 0.5$) leave our conclusions unchanged and are not reported to economise on space.[11]

Standardised values of the sorted observations are displayed in Figs. 2–4. According to the theory, observations displayed in the plots on the left

---

[10] Engel (1984) and Cumby (1988) find little difference in the behaviour of nominal deviations from uncovered interest parity and real deviations, suggesting that the Jensen's inequality problem is empirically unimportant.

[11] We use these parameter values to be as favourable as possible to the model. Cecchetti *et al.* (1994) found that the volatility restrictions implied by the intertemporal asset pricing model in pricing speculative currency returns on five USD forward foreign exchange rates could not be rejected when the preference parameters were set to these values. In addition, our results are quite robust. We made the computations under CRRA utility ($\delta = 0$) and durability ($\delta = 0.5$) with $\gamma = 57$. The large value of γ in these experiments is motivated by Kandel and Stambaugh (1990) who show that the equity premium puzzle and mean reversion in equity prices can be explained with $\gamma = 57$, while Bekaert (1994) finds that with $\gamma > 50$, the volatility of the intertemporal marginal rates of substitution satisfies the Hansen and Jagannathan (1991) volatility bounds implied by spot and forward exchange rate data. We also investigated the sensitivity of the calculations to Jensen's inequality by calculating results under CRRA utility under the assumption that the consumption growth and foreign exchange returns are log-normally distributed and found no substantive difference in the results.
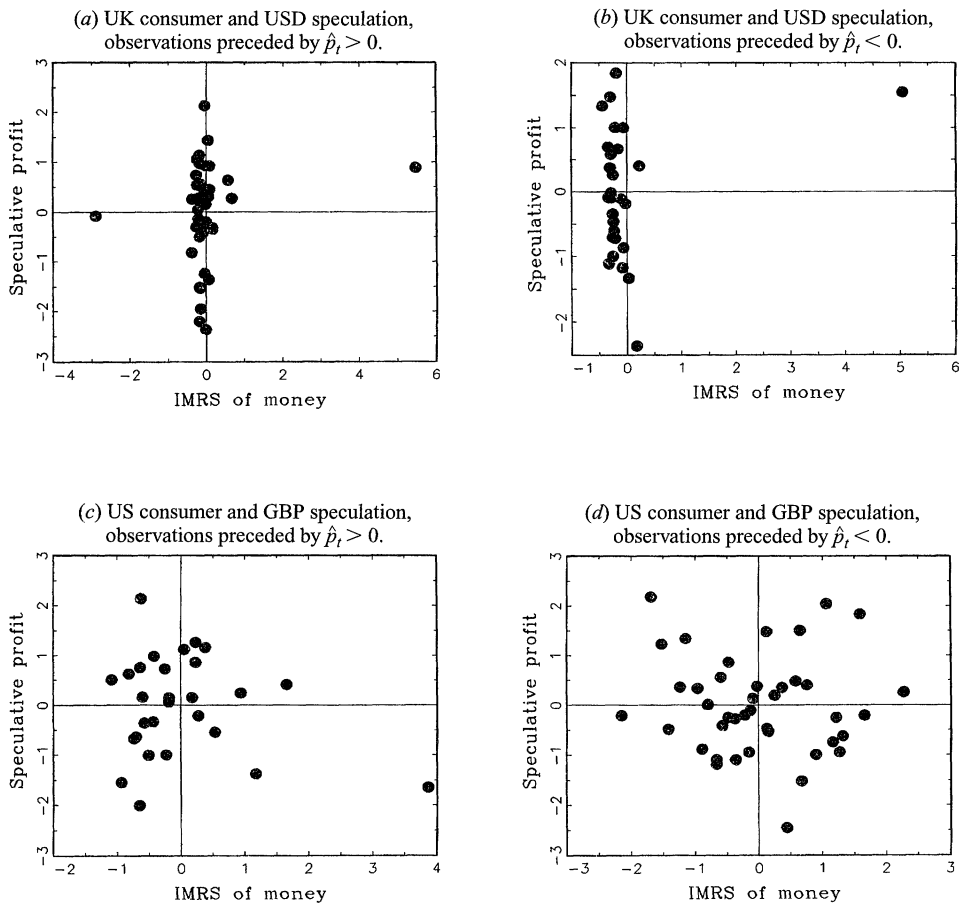
Fig. 2. *Scatter plots of standardised forward foreign exchange speculative profits $\tilde{s}_{t+1}$ and intertemporal marginal rate of substitution of money $m_{t+1} (\theta, \gamma, \delta)$ under habit persistence with $\theta = 0.99, \gamma = 7, \delta = -0.5$ sorted according to whether $\hat{p}_t$ is positive or negative.*

(Figs. $2a$, $2c$, $3a$, $3c$, $4a$, and $4c$) should be negatively correlated while observations displayed in the plots on the right should be positively correlated, but these predictions clearly are not borne out. As is evident, the data appear largely to be random.

We quantify these results by fitting regression lines through the scatter plots which we report in Table 2. Here, it is seen that the estimated slope coefficients typically do not have the sign predicted by the theory. When observations on returns and the intertemporal marginal rate of substitution are preceded by positive values of p_t, only in the case of the US individual who speculates on the GBP and DEM do the regression slope coefficients have the correct sign but the estimates are not statistically significant. When observations are preceded by negative values of p_t, we obtain the predicted sign in 2
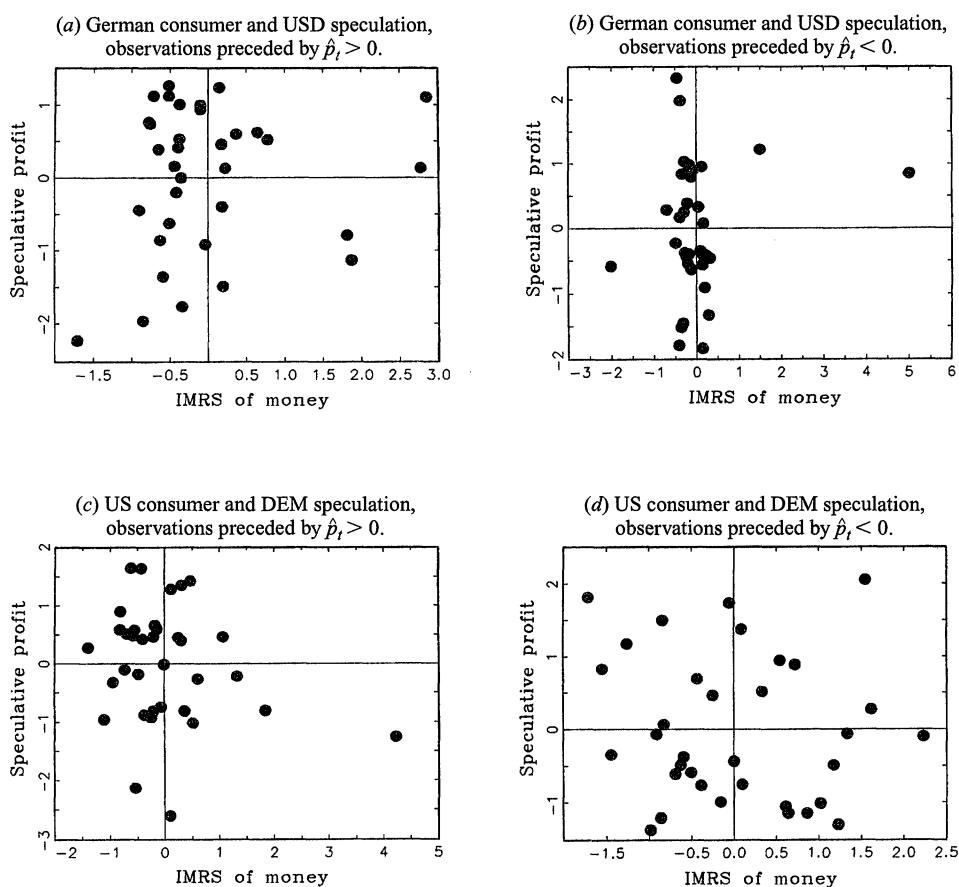
Fig. 3. *Scatter plots of standardised forward foreign exchange speculative profits $\tilde{s}_{t+1}$ and intertemporal marginal rate of substitution of money $m_{t+1}(\theta, \gamma, \delta)$ under habit persistence with $\theta = 0.99$, $\gamma = 7$, $\delta = -0.5$ sorted according to whether $\hat{p}_t$ is positive or negative.*

of the 6 cases, but again the estimates are not statistically significant. The inability of the standard model to produce a risk premium with the correct sign appears to pose a fundamental problem for the model in explaining the forward premium bias.[12]

3. Thinking About Noise

In this section, we study the overlapping-generations noise trader model of De Long *et al.* (1990) in the pricing of foreign currencies. Here, hetero-

[12] It is possible that the theory is true, but we have assumed and estimated the wrong data generating process in modelling $\hat{p}_t$. One possibility that the literature has examined is that occasional regime shifts create a 'peso problem.' However, Backus *et al.* (1994), Bekaert and Hodrick (1993), and Evans and Lewis (1995) show that peso problems alone cannot explain the entire magnitude of the forward premium bias.
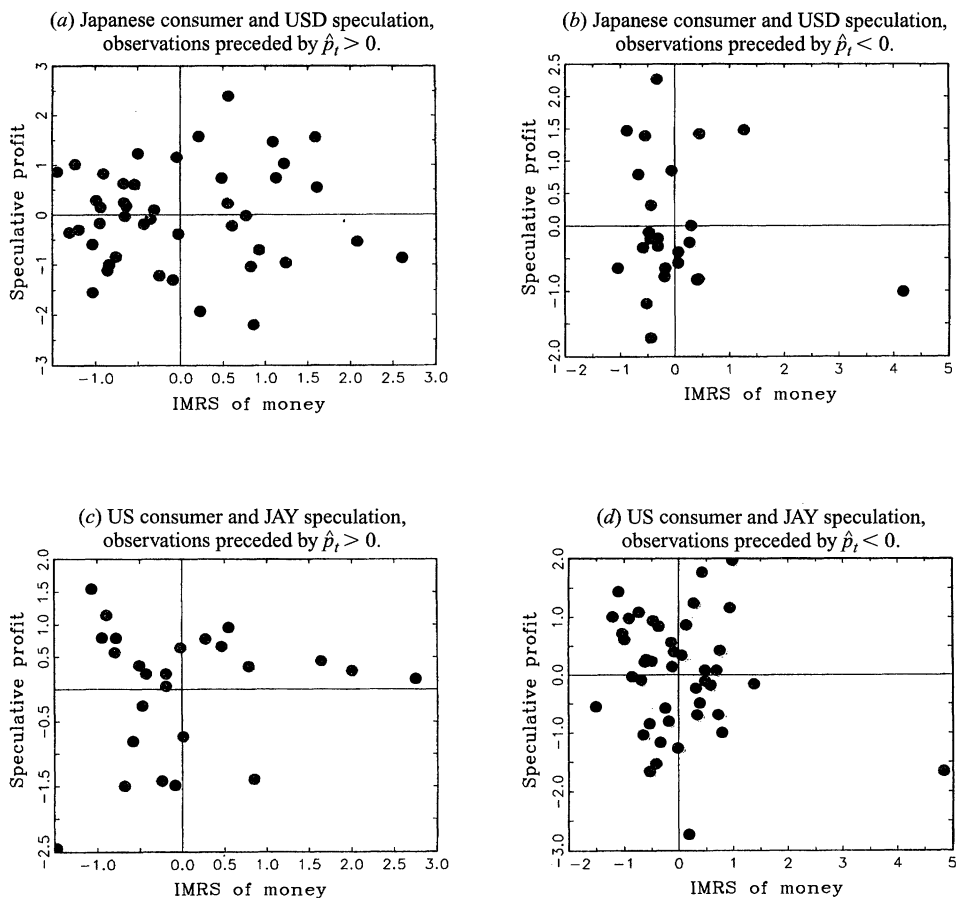
Fig. 4. *Scatter plots of standardised forward foreign exchange speculative profits* $\tilde{s}_{t+1}$ *and intertemporal marginal rate of substitution of money* $m_{t+1}(\theta, \gamma, \delta)$ *under habit persistence with* $\theta = 0.99, \gamma = 7, \delta = -0.5$ *sorted according to whether* $\hat{p}_t$ *is positive or negative.*

geneous beliefs across agents generate trading volume and excess currency returns. Black (1986) suggests that the environment of the real world is so complex that noise traders are unable to distinguish between pseudo-signals and news. Because these individuals think that the pseudo-signals contain information about economic fundamentals, their beliefs regarding prospective investment returns may be distorted by waves of excessive optimism and pessimism.[13] The resulting trading dynamics produce transitory deviations of the exchange rate from its fundamental value. Short-horizon rational investors bear the risk that they may be required to liquidate their positions at a

[13] De Bondt and Thaler (1986) report evidence of investor and financial analyst overreaction to news, while LeBaron (1992) and Taylor (1992) show that technical trading rules are at least as good as ARIMA models in predicting exchange rates.

Table 2

Regressions of Speculative Profits on the Intertemporal Marginal Rate of Substitution of Money Sorted According to Whether Observations are Preceded by Positive or Negative $\hat{p}_t$.

Investor	Currency	Observations preceded by $p_t > 0$			Observations preceded by $p_t < 0$		
		constant	slope	R^2	constant	slope	R^2
US	GBP	0.045	−0.039	0.026	0.003	−0.016	0.003
		(0.041)	(0.040)		(0.058)	(0.061)	
US	DEM	0.088	−0.072	0.142	0.011	−0.027	0.005
		(0.041)	(0.040)		(0.056)	(0.057)	
US	JAY	−0.012	0.035	0.013	0.029	−0.059	0.048
		(0.072)	(0.072)		(0.038)	(0.039)	
UK	USD	0.006	0.003	0.021	−0.014	0.003	0.029
		(0.011)	(0.003)		(0.011)	(0.003)	
German	USD	0.007	0.003	0.005	−0.037	0.019	0.075
		(0.013)	(0.007)		(0.019)	(0.015)	
Japanese	USD	0.011	0.012	0.007	−0.014	−0.012	0.023
		(0.023)	(0.022)		(0.025)	(0.018)	

Notes: Standard errors in parentheses. Consumption is measured by per capita expenditures on nondurables plus services. The sample extends from 1976:I–1994:I.

time when noise-traders have pushed asset prices even farther away from the fundamental value than they were when the investments were initiated.

3.1. *A Noise-Trader Model for Foreign Exchange*

We consider a two-country constant population partial equilibrium model. People are born with a full stomach, no assets, and live for two periods. The young do not consume but make portfolio decisions in order to maximise expected utility of second period wealth which is entirely consumed when old.

The home country currency unit is called the 'pound' and the foreign country currency unit is called the 'dollar.' In each country, there is a one-period asset that is safe in terms of the local currency. Both assets are available in perfectly elastic supply so that in period t, people can borrow or lend any amount they desire at the gross pound rate of interest R_t, or at the gross dollar rate of interest, R_t^*. The nominal interest differential – and hence by covered interest parity, the forward premium – is assumed to be exogenous. We base this assumption on the idea that interest rates reflect national economic conditions which are largely separate from currency movements while the forward rate is set simply to prevent covered interest arbitrage profits.

We assume there are legal restrictions on currency use. In order for financial wealth to have value, it must be denominated in the currency of the country in which the individual resides. Thus in the second period, the domestic agent converts wealth to pounds and the foreign agent converts wealth to dollars. We

also assume that the price level in each country is fixed at unity. Individuals therefore evaluate wealth in national currency units.

The domestic young decide whether to borrow pounds and lend dollars or *vice versa*. Let λ_t be the portfolio position taken. Long dollar positions are represented by positive values and long pound positions by negative values. To take a long dollar position, the young trader borrows λ_t pounds at the gross interest rate R_t and invests λ_t/S_t dollars at the gross rate R_t^*. When old, the dollar payoff $R_t^*(\lambda_t/S_t)$ is converted into $(S_{t+1}/S_t)R_t^*\lambda_t$ pounds. A long position in pounds is achieved by borrowing $-\lambda_t/S_t$ dollars and investing the proceeds in the pound asset at R_t. In the second period, the domestic agent sells $-(S_{t+1}/S_t)R_t^*\lambda_t$ pounds in order to repay the dollar debt $-R_t^*(\lambda_t/S_t)$. In either case, the net payoff is $[(S_{t+1}/S_t)R_t^* - R_t]\lambda_t$. We use the approximations $(S_{t+1}/S_t) \simeq (1 + \Delta s_{t+1})$ and $(R_t/R_t^*) = (F_t/S_t) \simeq 1 + x_t$ to express the net payoff as[14]

$$(\Delta s_{t+1} - x_t)R_t^*\lambda_t. \tag{7}$$

The foreign agent's portfolio position is denoted by λ_{*t} with positive values indicating long dollar positions. To take a long dollar position, the foreign young borrows λ_{*t} pounds and invests (λ_{*t}/S_t) dollars at the gross interest rate R_t^*. Next period's net dollar payoff is $(R_t^*/S_t - R_t/S_{t+1})\lambda_{*t}$. A long pound position is achieved by borrowing $-(\lambda_{*t}/S_t)$ dollars and investing $-\lambda_{*t}$ pounds. The net dollar payoff in the second period is $-(R_t/S_{t+1} - R_t^*/S_t)\lambda_{*t}$. Using the approximation $(F_t S_t)/(S_t S_{t+1}) \simeq 1 + x_t - \Delta s_{t+1}$, we express the net dollar payoff as

$$(\Delta s_{t+1} - x_t)R_t^*\frac{\lambda_{*t}}{S_t}. \tag{8}$$

The foreign exchange market clears when net pound sales of the current young equals net pound purchases of the current old,

$$\lambda_t + \lambda_{*t} = \frac{S_t}{S_{t-1}}R_{t-1}^*\lambda_{t-1} + R_{t-1}\lambda_{*t-1}. \tag{9}$$

Fundamental and Noise Traders. A fraction μ of domestic and foreign traders are fundamentalists who have rational expectations. The remaining fraction $1 - \mu$ are noise traders whose beliefs concerning future returns from their portfolio investments are distorted. Let the speculative positions of home fundamentalist and home noise traders be given by λ_t^f and λ_t^n respectively. Similarly, let foreign fundamentalist and foreign noise trader positions be λ_{*t}^f and λ_{*t}^n. The total portfolio position of domestic residents is $\lambda_t = \mu\lambda_t^f + (1 - \mu)\lambda_t^n$ and of foreign residents is $\lambda_{*t} = \mu\lambda_{*t}^f + (1 - \mu)\lambda_{*t}^n$.

We denote subjective date t conditional expectations generically as $\mathscr{E}_t(\cdot)$. When the distinction is necessary, expectations of fundamentalists will be

---

[14] These approximations are necessary in order to avoid dealing with Jensen inequality terms when evaluating the foreign wealth position which render the model intractable.

denoted by $E_t(\cdot)$. Similarly, the conditional variance is generically denoted by $\mathscr{V}_t(\cdot)$ with the conditional variance of fundamentalists denoted by $V_t(\cdot)$.

Utility displays CARA (constant absolute risk aversion) with coefficient γ. The young construct a portfolio to maximise the expected utility of next period wealth,

$$\mathscr{E}_t(-e^{-\gamma W_{t+1}}). \tag{10}$$

Both fundamental and noise traders believe that conditional on time t information, W_{t+1} is normally distributed. Since (10) is the (negative of) the conditional moment generating function of W_{t+1}, maximising (10) under normality is equivalent to maximising

$$\mathscr{E}_t(W_{t+1}) - \frac{\gamma}{2} \mathscr{V}_t(W_{t+1}). \tag{11}$$

A Fundamentals $(\mu = 1)$ *Economy.* We begin by assuming that everyone is rational $(\mu = 1)$ so that $\mathscr{E}_t(\cdot) = E_t(\cdot)$ and $\mathscr{V}_t(\cdot) = V_t(\cdot)$. Total second period wealth of the fundamentalist domestic agent is the portfolio payoff plus c pounds of exogenous 'labour' income which is paid in the second period.[15] We simplify the exposition by assuming that R^* is fixed at unity. The forward premium, $(F_t/S_t) = (R_t/R^*) = R_t \simeq 1 + x_t$ thus inherits its stochastic properties from R_t. Accordingly, we assume,

$$x_t = \rho x_{t-1} + v_t, \tag{12}$$

where $0 < \rho < 1$, and $v_t \sim$ i.i.d. with mean 0 and variance σ_v^2. Second period wealth can now be written as

$$W_{t+1}^f = (\Delta s_{t+1} - x_t)\lambda_t^f + c. \tag{13}$$

People evaluate the conditional mean and variance of next period wealth as

$$E_t(W_{t+1}^f) = [E_t(\Delta s_{t+1}) - x_t]\lambda_t^f + c, \tag{14}$$

$$V_t(W_{t+1}^f) = \sigma_s^2(\lambda_t^f)^2, \tag{15}$$

where $\sigma_s^2 = V_t(\Delta s_{t+1})$.[16] The domestic fundamental trader's problem is to choose λ_t^f to maximise

$$[E_t(\Delta s_{t+1}) - x_t]\lambda_t^f + c - \frac{\gamma}{2}(\lambda_t^f)^2\sigma_s^2, \tag{16}$$

which is attained by setting

$$\lambda_t^f = \frac{[E_t(\Delta s_{t+1}) - x_t]}{\gamma\sigma_s^2}. \tag{17}$$

---

[15] The exogenous income is introduced to lessen the likelihood of negative second period wealth realisations, but as in De Long *et al.*, we cannot rule out such a possibility.

[16] Baillie and Bollerslev (1989) find little evidence that percentage changes in nominal exchange rates are conditionally heteroscedastic beyond the 1-week horizon. Accordingly, we assume that $\{\Delta s_t\}$ is a conditionally homoscedastic process with mean zero and fixed variance σ_s^2.

Equation (17) displays the familiar property of CARA utility in which portfolio positions are proportional to the expected asset payoff. The factor of proportionality is inversely related to the individual's absolute risk aversion coefficient. Recall that individuals undertake zero-net investment strategies. The portfolio position in our setup does not depend on wealth because traders are endowed with zero initial wealth.

The foreign fundamental trader faces an analogous problem. The second period dollar-wealth of fundamentalist foreign agents is the payoff from portfolio investments plus an exogenous dollar payment of 'labour' income c_*, $W^f_{*\,t+1} = (\Delta s_{t+1} - x_t)(\lambda^f_{*\,t}/S_t) + c_*$. The solution is to choose $\lambda^f_{*\,t} = S_t\lambda^f_t$. Because individuals at home and abroad have identical tastes but evaluate wealth in national currency units, they will pursue identical investment strategies by taking positions of the same size as measured in monetary units of the country of residence.

These portfolios combined with the market clearing condition (9) imply the difference equation,

$$E_t\Delta s_{t+1} - x_t = \Gamma_t(E_{t-1}\Delta s_t - x_{t-1}),\qquad(18)$$

where $\Gamma_t \equiv [(S_t/S_{t-1}) + S_{t-1}R_{t-1}]/(1 + S_t)$. The level of the exchange rate is indeterminate but it is easily seen that a solution for the rate of depreciation is

$$\Delta s_t = \frac{1}{\rho}x_t = x_{t-1} + \frac{1}{\rho}v_t.\qquad(19)$$

The independence of v_t and x_{t-1} implies $E_t(\Delta s_{t+1}) = x_t$ and the fundamentals solution therefore does not generate a forward premium bias. Because the exchange rate is uncorrelated with other prices in the model, exchange rate risk is diversifiable just as it is in Frankel's (1979) setup with no outside bonds. Since the market does not compensate investors for diversifiable risk, the model does not produce a risk premium.

A Noise Trader ($\mu < 1$) Economy. We now introduce noise traders whose beliefs about expected returns are distorted by the stochastic process $\{n_t\}$. Noise traders are capable of computing $E_t(x_{t+1})$, but believe that factors in addition to news affect returns. The distortion in noise trader beliefs occurs only in evaluating first moments of returns. Their evaluation of second moments coincide with those of fundamentalists. The current young domestic noise trader evaluates the conditional mean and variance of next period wealth as

$$\mathscr{E}_t(W^n_{t+1}) = [E_t(\Delta s_{t+1}) - x_t]\lambda^n_t + n_t\lambda^n_t + c,\qquad(20)$$

$$\mathscr{V}_t(W^n_{t+1}) = (\lambda^n_t)^2\sigma^2_s.\qquad(21)$$

Recall that a positive value of λ_t represents a long position in dollars. Equation (20) implies that noise traders will appear to overreact to news and to display excess pound pessimism when $n_t > 0$ for they believe the pound will be weaker in the future than is justified by the fundamentals.

To complete the specification of the model, we draw on results from Froot and Frankel's (1989) empirical study of survey expectations in which they find that respondents place excessive weight on the forward premium when predicting future changes in the exchange rate. We build upon this idea and parameterise the noise trader distortion as

$$n_t = kx_t + u_t, \tag{22}$$

where $k > 0$, $\{u_t\} \sim$ i.i.d. with $E(u_t) = 0$ and $Var(u_t) = \sigma_u^2$. The domestic noise trader's problem is to maximise $\lambda_t^n (E_t \Delta s_{t+1} - x_t + n_t) - \gamma (\lambda_t^n)^2 \sigma_s^2 / 2$. The solution is to choose

$$\lambda_t^n = \lambda_t^f + \frac{n_t}{\gamma \sigma_s^2}, \tag{23}$$

from which it can be seen that the noise trader's position deviates from that of the fundamentalist by a term that depends on the distortion in their beliefs, n_t.

The foreign noise trader holds similar beliefs, solves an analogous problem and chooses

$$\lambda_{*t}^n = S_t \lambda_t^n. \tag{24}$$

Substituting these optimal portfolio positions into the market clearing condition (9) yields the stochastic difference equation

$$(E_t \Delta s_{t+1} - x_t) + (1 - \mu) n_t = \Gamma_t [(E_{t-1} \Delta s_t - x_{t-1}) + (1 - \mu) n_{t-1}]. \tag{25}$$

Using the method of undetermined coefficients, it can be verified that

$$\Delta s_t = \frac{1}{\rho} x_t - \frac{(1 - \mu)}{\rho} n_t - (1 - \mu) u_{t-1}, \tag{26}$$

is a solution.

Properties of the solution. First, fundamentalists and noise traders both believe, *ex ante*, that they will earn positive profits from their portfolio investments. It is the differences in their beliefs that lead them to take opposite sides of the transaction. When noise traders are excessively pessimistic and take short positions in the pound, fundamentalists take the offsetting long position. In equilibrium, the expected payoff of fundamentalists and noise-traders are respectively,

$$E_t \Delta s_{t+1} - x_t = -(1 - \mu) n_t, \tag{27}$$

$$\mathscr{E}_t \Delta s_{t+1} - x_t = \mu n_t. \tag{28}$$

On average, the forward premium is the subjective predictor of the future depreciation: $\mu E_t \Delta s_{t+1} + (1 - \mu) \mathscr{E}_t \Delta s_{t+1} = x_t$. As the measure of noise traders approaches 0 ($\mu \to 1$), the fundamentals solution with no trading is restored. Foreign exchange risk, excess currency movements, and trading

volume are induced entirely by noise traders. Neither type of trader is guaranteed to earn profits or losses, however. The *ex post* profit depends on the sign of

$$\Delta s_{t+1} - x_t = -(1 - \mu)\, n_t + \frac{1}{\rho}[1 - k(1 - \mu)]v_{t+1} - \frac{1 - \mu}{\rho}\, u_{t+1}, \qquad (29)$$

which can be positive or negative.

Second, the model is able to generate a negative forward premium bias. Substituting expressions (22) and (12) into (26) yields,

$$\Delta s_{t+1} = [1 - k(1 - \mu)]x_t + \xi_{t+1}, \qquad (30)$$

where $\xi_{t+1} \equiv (1/\rho)[1 - k(1 - \mu)]v_{t+1} - (1 - \mu)/\rho u_{t+1} - (1 - \mu)u_t$ is orthogonal to x_t. The implied slope coefficient in a regression of the future depreciation on the forward premium will be negative provided that $1 - k(1 - \mu) < 0$. If we assume it is, then it is also true that p_t covaries negatively with and is more volatile than the rationally expected depreciation, $E_t \Delta s_{t+1}$. This can be seen from the implied second moments of these two random variables which are,

$$\text{Cov}\{[x_t - E_t(\Delta s_{t+1})], E_t(\Delta s_{t+1})\} = k(1 - \mu)[1 - k(1 - \mu)]\sigma_x^2 - (1 - \mu)^2\sigma_u^2,$$

$$(31)$$

$$\text{Var}[x_t - E_t(\Delta s_{t+1})] = (1 - \mu)^2 (k^2\sigma_x^2 + \sigma_u^2), \qquad (32)$$

$$\text{Var}[E_t(\Delta s_{t+1})] = \text{Var}[x_t - E_t(\Delta s_{t+1})] + [1 - 2k(1 - \mu)]\sigma_x^2.$$

$$(33)$$

3.2. *Evidence from Survey Expectations*

Empirical research on the properties of surveys of foreign exchange forecasts by professional market participants provides fragmentary evidence favorable to the noise–trader model. We draw on three studies in the literature that employ different data sets over separate time periods. The first is Froot and Frankel (1989), who study surveys conducted by the Economist (*Economist's Financial Report*) from 6/81–12/85, MMS (Money Market Services) from 1/83–10/84, and AMEX (American Express Banking Corporation) from 1/76–7/85. The second study that we draw on is Frankel and Chinn (1993) who employ a survey compiled monthly by CFD (*Currency Forecasters' Digest*) from 2/88 to 2/91. The third study is by Cavaglia *et al.* (1994) who analyse forecasts on 10 USD bilateral rates and 8 DEM bilateral rates surveyed by BIC (Business International Corporation) from 1/86 to 12/90. The survey respondents were asked to provide forecasts at horizons of 3, 6, and 12 months into the future.

Let $\Delta \hat{s}^e_{t+1}$ denote the forecasted 1–period ahead percent change in the exchange rate reported by survey respondents. Equation (2) and the following two regressions are typically estimated in these studies:

$$\Delta \hat{s}^e_{t+1} - \Delta s_{t+1} = \alpha_1 + \beta_1 x_t + \epsilon_{1,t+1}, \tag{34}$$

$$\Delta \hat{s}^e_{t+1} = \alpha_2 + \beta_2 x_t + \epsilon_{2,t+1}. \tag{35}$$

Froot and Frankel and Cavaglia *et al.* estimate (2), (34) and (35) while Frankel and Chinn estimate (2) and (35). It can be seen that the three equations are not independent and that $\beta = \beta_2 - \beta_1$. Froot and Frankel and Cavaglia *et al.* test the hypothesis that survey respondents are rational ($\beta_1 = 0$) while all three studies test the hypothesis that there is no time-varying risk premium ($\beta_2 = 1$). In Table 3, we display Froot and Frankel's pooled estimates for the *Economist*, MMS, and AMEX data sets and Frankel and Chinn's fixed-effects slope estimates for the CFD. Cavaglia *et al.* report a massive number of estimates on individual bilateral exchange rates but do not report pooled estimates. We summarise their results by taking the average of their 10 USD exchange rate estimates (BIC-Dollar) and the average of their 8 DEM exchange rate estimates (BIC-DEM). We note that the forward premium bias is present during the time periods in which each of the surveys are conducted as the summary estimates of β are negative in each case. All three studies conclude that the survey forecasts exhibit signs of irrationality. The form in which rationality is violated ($\beta_1 > 0$) suggests that actual participants place excess weight on the forward premium which was the motivation for our specification of noise-trader over-

Table 3
Empirical Estimates from Studies of Survey Forecasts

	Data Set					
	Economist	MMS	AMEX	CFD	BIC–USD	BIC–DEM
	(a) Horizon: 3-months					
β	−1.209	−6.254		−2.881	−4.028	−0.971
β_1	2.513	6.073			5.971	1.930
β_2	1.304	−0.182		0.423	1.930	0.959
t-*statistic*, H_0: $\beta_2 = 1$	1.188	−2.753		−2.842	5.226	−1.452
	(b) Horizon: 6-months					
β	−1.982		−2.418		−3.857	−1.026
β_1	2.986		3.635		5.347	1.841
β_2	1.033		1.216		1.222	0.812
t-*statistic*, H_0: $\beta_2 = 1$	0.192		1.038		1.461	−4.325
	(c) Horizon: 12-months					
β	0.289		−2.138	−3.409	−4.554	−1.203
β_1	0.517		3.108		5.601	1.706
β_2	0.929		0.877	1.055	1.046	0.502
t-*statistic*, H_0: $\beta_2 = 1$	−0.476		−0.446	0.297	0.532	−6.594

Notes: Estimates from the *Economist*, MMS, and AMEX surveys are from Froot and Frankel (1989). Estimates from the CFD survey are from Frankel and Chinn (1993), and estimates from BIC survey from Cavaglia *et al.* (1994). BIC–USD is the average of individual estimates for 10 USD exchange rates. BIC–DEM is the average over 8 DEM exchange rates. β is the slope coefficient from the regression of the depreciation on the forward premium over the survey period, β_1 is the slope coefficient from the regression of the survey forecast error on the forward premium, and β_2 is the slope coefficient from the regression of the survey forecast on the forward premium.

reaction. Cavaglia *et al.* and Frankel and Chinn find some evidence that there is a time-varying risk premium while Froot and Frankel do not.

To bring the research on survey expectations to bear on the noise-trader model, suppose that the survey sampled fundamentalists and noise traders in the same proportion that they exist in population. Then $\Delta \hat{s}^e_{t+1} = \mu E_t(\Delta s_{t+1}) + (1-\mu)\mathscr{E}_t(\Delta s_{t+1})$ which implies $\beta_2 = 1$.

Froot and Frankel's and Frankel and Chinn's results allow valid t-tests of the hypothesis $\beta_2 = 1$, which is generally not rejected at the 5% level. The results of Cavaglia *et al.* are interpreted as follows: let t_i be the t-statistic for the individual $i = 1, \ldots, N$ test. We formulate a quasi t-test by calculating $(1/\sqrt{N})\sum t_i$ which is asymptotically standard normal if the t_i are independent. The tests conducted under the independence assumption yield some evidence against the hypothesis that $\beta_2 = 1$ only for the USD at the 3 month horizon, and for the DEM at 6 and 12 months.

While the above calculations yield little evidence against the noise-trader model, our assumption that the proportion of noise traders sampled in the survey is equal to that in the population makes it impossible to identify k and μ. However, if it is assumed that the share of fundamentalist traders in the survey, m, measures the population proportions with error, the survey expected depreciation becomes, $\Delta \hat{s}^e_{t+1} = m E_t(\Delta s_{t+1}) + (1-m)\mathscr{E}_t(\Delta s_{t+1})$. For a given m, the model implies $k = \beta_1/(1-m)$ and $\mu = (\beta - 1 + k)/k$.

We show implied values of k and μ in Table 4 for $m = 0.3, 0.5, 0.7$. The implied values of μ are generally close to the assumed m values. Larger values of m result in larger values of k and smaller values of μ. For a given value of the forward premium, $k\mu$ measures the excess depreciation anticipated by noise traders beyond what is implied by the forward premium. For example, if half of the survey respondents are fundamentalists ($m = 0.5$) and the forward premium is zero, over the next 12 months, noise traders expect a depreciation of 0.32% in the *Economist* survey, 3.10% in the AMEX survey, 4.52% in the CFD survey, 5.65% in the BIC-Dollar survey, and 1.21% in the BIC-DEM survey. $m = 0.7$ implies values of k and noise-trader over-reaction that may be too large to be reasonable. Taking $m = 0.3$, on the other hand, generates plausible values of the parameters.

4. Conclusion

Our examination of the standard rational intertemporal asset pricing model suggests that it has little empirical content in interpreting the forward premium bias. While the model provides an intuitive and appealling theory of the forward foreign exchange risk premium, one of the model's sharpest implications – that the sign of the deviation from rational uncovered interest parity is determined by the sign of the conditional covariance between the intertemporal marginal rate of substitution of money and the payoff from forward speculation – lacks empirical support. The problems with the model evidently lie at a more primitive level than an inability to produce sufficiently volatile returns, which is where the literature has recently focused.

Table 4

Implied Parameter Values from Studies of Survey Forecasts

Horizon	Implied Values	Data Set					
		Economist	MMS	AMEX	CFD	BIC–USD	BIC–DEM
		(a) $m = 0.3$					
3	k	3.590	8.676		4.720	8.530	2.758
	μ	0.385	0.164		0.178	0.410	0.285
	$k\mu$	1.382	1.423		0.839	3.497	0.786
6	k	4.266		5.192		7.638	2.631
	μ	0.301		0.342		0.364	0.230
	$k\mu$	1.284		1.776		2.780	0.605
12	k	0.739		4.440	6.377	8.002	2.437
	μ	0.037		0.293	0.309	0.306	0.096
	$k\mu$	0.027		1.301	1.968	2.449	0.234
		(b) $m = 0.5$					
3	k	5.026	12.146		6.608	11.942	3.861
	μ	0.560	0.403		0.413	0.579	0.489
	$k\mu$	2.815	4.895		2.727	6.914	1.888
6	k	5.972		7.269		10.694	3.683
	μ	0.501		0.530		0.546	0.450
	$k\mu$	2.992		3.853		5.839	1.657
12	k	1.034		6.216	8.928	11.203	3.412
	μ	0.312		0.495	0.506	0.504	0.354
	$k\mu$	0.323		3.077	4.519	5.646	1.208
		(c) $m = 0.7$					
3	k	8.377	20.243		11.013	19.903	6.434
	μ	0.736	0.642		0.648	0.747	0.694
	$k\mu$	6.156	12.996		7.132	14.868	4.465
6	k	9.953		12.115		17.823	6.138
	μ	0.700		0.718		0.727	0.670
	$k\mu$	6.967		8.699		12.957	4.112
12	k	1.723		10.360	14.880	18.671	5.686
	μ	0.587		0.697	0.704	0.703	0.612
	$k\mu$	1.011		7.221	10.470	13.126	3.480

Notes: See notes to Table 3 for definitions. m is the fraction of the survey respondents assumed to be fundamentalists.

While the search for a rational theory of the forward premium bias has proved elusive, a potentially promising alternative approach is the quasi-rational noise–trader model of De Long *et al.*. In addition to providing an account of the forward premium bias, the model supplies an explanation for foreign exchange trading, and for transient deviations of the exchange rate from its fundamental value. Although the model is highly stylised and cannot reasonably be expected to match all dimensions of the data, estimates from the literature on survey expectations provide fragmentary support.

Ohio State University
Rutgers University

References

Backus, D., Gregory, A. and Telmer, C. (1993). 'Accounting for forward rates in markets for foreign currency,' *Journal of Finance*, vol. 48, (December), pp. 1887–908.

Backus, D., Foresi, S. and Telmer, C. (1994). 'The forward premium anomaly: three examples in search of a solution,' Columbia PaineWebber Working Paper Series in Money, Economics and Finance: 95-01, (July).

Baillie, R. and Bollerslev, T. (1989). 'The message in daily exchange rates: a conditional-variance tale,' *Journal of Business and Economic Statistics*, vol. 7, (July), pp. 297–305.

Bekaert, G. (1994). 'Exchange rate volatility and deviations from unbiasedness in a cash-in-advance model,' *Journal of International Economics*, vol. 36, (February), pp. 29–52.

Bekaert, G. and Hodrick, R. (1992). 'Characterizing predictable components in excess returns on equity and foreign exchange markets,' *Journal of Finance*, vol. 48, (June), pp. 467–509.

Bekaert, G. and Hodrick, R. (1993). 'On biases in the measurement of foreign exchange risk premiums,' *Journal of International Money and Finance*, vol. 12, (April), pp. 115–38.

Breuer, J. and Wohar, M. (1996). 'The road less travelled: institutional aspects of data and their influence on empirical estimates with an application to tests of forward rate unbiasedness,' ECONOMIC JOURNAL, vol. 106, (January), pp. 26–38.

Black, F. (1986). 'Noise,' *Journal of Finance*, vol. 41, (July), pp. 529–43.

Cavaglia, S., Verschoor, W. and Wolff, C. (1994). 'On the biasedness of forward foreign exchange rates: irrationality or risk premia?' *Journal of Business*, vol. 67, (July), pp. 321–43.

Canova, F. and Ito, T. (1991). 'The time-series properties of the risk premium in the yen/dollar exchange market,' *Journal of Applied Econometrics*, vol. 6, (April–June), pp. 125–42.

Cecchetti, S., Lam, P. and Mark, N. (1994). 'Testing volatility restrictions on intertemporal marginal rates of substitution implied by euler equations and asset returns,' *Journal of Finance*, vol. 49, (March), pp. 123–52.

Chen, J. and Mark, N. (1996). 'Alternative long-horizon exchange rate predictors,' *International Journal of Finance and Economics*, vol. 1, (October), pp. 229–50.

Cheung, Y. (1993). 'Exchange rate risk premiums,' *Journal of International Money and Finance*, vol. 12, (April), pp. 182–94.

Chinn, M. and Meese, R. (1995). 'Banking on currency forecasts: how predictable is change in money?' *Journal of International Economics*, vol. 38, (February), pp. 161–78.

Cumby, R. (1988). 'Is it risk? explaining deviations from uncovered interest parity,' *Journal of Monetary Economics*, vol. 22, (September), pp. 279–99.

De Bondt, W. and Thaler, R. (1986). 'Does the stock market overreact?,' *Journal of Finance*, vol. 40, (July), pp. 793–807.

De Long, B., Shleifer, A., Summers, L. and Waldman, R. (1990). 'Noise trader risk in financial markets,' *Journal of Political Economy*, vol. 98, (August), pp. 703–38.

Eichengreen, B., Tobin, J. and Wyplosz, C. (1995). 'Two cases for sand in the wheels of international finance,' ECONOMIC JOURNAL, vol. 105, (January), pp. 162–72.

Engel, C. (1996). 'Why is there a forward discount bias? a survey of recent evidence,' *Journal of Empirical Finance*, vol. 3, (June), pp. 123–91.

Engel, C. (1984). 'Testing for the absence of expected real profits from forward market speculation,' *Journal of International Economics*, vol. 17, (November), pp. 309–24.

Evans, M. and Lewis, K. (1993). 'Trends in excess returns in currency and bond markets,' *European Economic Review*, vol. 37, (June), pp. 1005–19.

Evans, M. and Lewis, K. (1995). 'Do long-term swings in the dollar affect estimates of the risk premia?' *Review of Financial Studies*, vol. 8, (Fall), pp. 709–42.

Fama, E. (1984). 'Forward and spot exchange rates,' *Journal of Monetary Economics*, vol. 14, (November), pp. 319–38.

Frankel, J. (1979). 'The diversifiability of exchange risk,' *Journal of International Economics*, vol. 9, (August), pp. 379–93.

Frankel, J. and Chinn, M. (1993). 'Exchange rate expectations and the risk premium: tests for a cross section of 17 currencies,' *Review of International Economics*, vol. 1, (June), pp. 136–44.

Froot, K. and Frankel, J. (1989). 'Forward discount bias: is it an exchange risk premium?' *Quarterly Journal of Economics*, vol. 104, (February), pp. 139–61.

Froot, K. and Thaler, R. (1990). 'Anomalies: foreign exchange,' *Journal of Economic Perspectives*, vol. 4, (Summer), pp. 179–92.

Goldberg, M. and Frydman, R. (1996). 'Imperfect knowledge and behaviour in the foreign exchange market,' ECONOMIC JOURNAL, vol. 106, (July), pp. 869–93.

Hai, W., Mark, N. and Wu, Y. (1997). 'Understanding spot and forward exchange rate regressions,' *Journal of Applied Econometrics*, vol. 12, (November–December), pp. 715–34.

Hansen, L. and Hodrick, R. (1983). 'Risk averse speculation in the forward foreign exchange market: an econometric analysis of linear models,' In *Exchange Rates and International Macroeconomics*, (ed.

J. Frenkel,) Chicago: University of Chicago Press for NBER.

Hansen, L. and Jagannathan, R. (1991). 'Implications of security market data for models of dynamic economies,' *Journal of Political Economy*, vol. 91, (April), pp. 225–62.

Hodrick, R. (1987). *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets*, Chur:Harwood.

Kandel, S. and Stambaugh, R. (1990). 'Expectations and volatility of consumption and asset returns,' *Review of Financial Studies*, vol. 3, pp. 207–32.

Lewis, K. (1995). 'Puzzles in international financial markets,' In *Handbook of International Economics* (ed. G. Grossman and K. Rogoff), vol. 3, pp. 1914–71, North Holland: Amsterdam.

Lothian J. and Taylor, M. (1996). 'Real exchange rate behavior: the recent float from the perspective of the past two centuries,' *Journal of Political Economy*, vol. 104, (June), pp. 488–509.

LeBaron, B. (1992). 'Do moving average trading rule results imply nonlinearities in foreign exchange markets?' University of Wisconsin SSRI working paper no. 9222.

Mark, N. (1995). 'Exchange rates and fundamentals: evidence on long-horizon predictability,' *American Economic Review*, vol. 85, (March), pp. 201–18.

Mark, N. and Choi, D. (1997). 'Real exchange-rate prediction over long horizons,' *Journal of International Economics*, vol. 43, (August), pp. 29–60.

Taylor, S. (1992). 'Rewards available to currency futures speculators: compensation for risk or evidence of inefficient pricing?' *Economic Record*, vol. 68 Supplement, pp. 105–16.

Wolff, C. (1987). 'Forward foreign exchange rates, expected spot rates, and premia: a signal–extraction approach,' *Journal of Finance*, vol. 42, (June), pp. 295–406.

Wu, Y. and Zhang, H. (1997). 'Forward premiums as unbiased predictors of future currency depreciation: a non-parametric approach,' *Journal of International Money and Finance*, vol. 16, (August), pp. 609–23.

Appendix

Exchange rates. Spot and 3-month forward exchange rates are taken from Harris Bank's *Weekly Review*. They are drawn from those Fridays occurring nearest to the end of the calendar quarter.

Per capita consumption. For the United States, we use consumption expenditure on nondurables plus services divided by population. Both consumption and population data are from Citibank's *Citibase* data bank. For the other three countries, aggregate consumption expenditure and population data are taken from IMF's *International Financial Statistics* (*IFS*). Prices for all countries are measured by consumer price indices, which also were taken from the *IFS*.