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A Note on the Foreign Exchange Market Efficiency Hypothesis: Does Small Sample Bias affect Inference?

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Key Words: Market Efficiency Hypothesis; Rational Expectation Hypothesis; Risk Premium, Small Sample Bias; Bootstrapping

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A Note on the Foreign Exchange Market Efficiency Hypothesis: Does Small Sample Bias affect Inference?

Abstract

This study examines whether small sample bias affects the standard inference about the foreign exchange market efficiency hypothesis. Our findings indicate that the bias is large enough to result in rejection of the efficient market hypothesis even when it is true. We use bootstrapping to adjust for the bias and find that the hypothesis cannot be rejected for the Swiss franc and French franc. We also find that the bias plays a significant role in the inference that expectation error causes inefficiency in the foreign exchange markets. After bias adjustment, the rational expectation hypothesis holds even at one month-horizon.

I. Introduction

There is a large literature that studies the efficiency of the foreign exchange markets and its implications for the market participants. The traditional tests of the foreign exchange market efficiency hypothesis, hereafter MEH, are based on a linear projection of the forward rate on the future spot exchange rate.[1] To circumvent the non-stationarity problem in this estimation procedure, Froot and Frankel (1989) use the forward premium as the regressor and the exchange rate differential as the regressand. As shown in Liu and Maddala (1992), this adjustment can lead to inconsistent estimate of the slope coefficient because the forward rate is correlated with the risk premium. Liu and Maddala suggest therefore regressing the forward premium on the exchange rate differential when both series are stationary. We note, however, that this approach can exhibit finite sample bias due to the presence of an endogenous regressor. Whether the small sample bias is large enough to result in rejection of the MEH even when it is true remains to our knowledge an open empirical question.

In this paper, we examine whether small sample bias plays a role in the inference about the MEH. We employ bootstrapping technique to obtain bias adjusted estimates and confidence intervals. To account for the leptokurtic distribution of exchange rates, we use the Least Absolute Deviation estimation method. We also investigate the impact of small sample bias on the rational expectation hypothesis (REH), which is one of the explanations for the failure of the MEH. To test these two hypotheses, we use monthly exchange rate data on the pound sterling, Deutsche mark, Japanese yen, Swiss franc, and French franc.

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[1] See, for example, Frenkel (1977), Longworth (1981), and Jacobs (1982).

The results indicate that the MEH is rejected too often because of the presence of small sample bias. After adjusting for this bias, we find that the MEH cannot be rejected for the Swiss franc and French franc. The role of small sample bias is even more significant for the REH. Across all currencies, the REH cannot be rejected after we have corrected for the small sample bias. This result is in contrast with the findings in Liu and Maddala (1992). Using standard methods, they find that the REH is rejected for monthly data.

Consistent with Liu and Maddala (1992) and Naka and Whitney (1995) we find evidence supporting the existence of foreign exchange rate risk premium. The Breitung (2002) stationarity tests suggest however that this risk premium is not present in all the foreign exchange markets. We also find that the rejection of the MEH for the pound sterling, Deutsche mark, and Japanese yen is not due to expectation errors, but rather to the existence of a risk premium.

The remainder of the paper is organized as follows. In Section II, we specify and explain the small sample bias in standard tests of the market efficiency hypothesis. Section III describes the methodology and data. In Section IV, we discuss the estimation and test results. Section V concludes.

II. Small Sample Bias

The foreign exchange market efficiency hypothesis (MEH) contends that the forward exchange rate, F_t, is an unbiased predictor of the future spot exchange rate, S_{t+1}. Several studies on the MEH are based on regressing S_{t+1} on F_t ,

$$S_{t+1} = \alpha + \beta F_t + \varepsilon_{t+1}, \tag{1}$$

where H_0: $\alpha = 0$, $\beta = 1$. As is well known, if S_t and F_t are unit root, but not cointegrated processes, then standard tests can lead to incorrect inference. To avoid this non-stationarity problem, Froot and Frankel (1989) adjust equation (1) by using differences instead of levels,

$$\left(S_{t+1} - S_t\right) = \alpha + \beta\left(F_t - S_t\right) + \varepsilon_{t+1}. \tag{2}$$

They find that the rejection of the MEH can be attributed primarily to the failure of the rational expectation hypothesis (REH).

Liu and Maddala (1992) show that the estimation of equation (2) can give inconsistent estimate of β in the presence of foreign exchange risk premium and can lead to rejection of the MEH even when it is true. More specifically, they show that if S_t and F_t are $I(1)$ processes and both (S_{t+1} - S_t) and (F_t - S_t) are stationary, then OLS estimation is inconsistent because the risk premium is correlated with the forward rate. To remedy this inconsistency problem in equation (2), Liu and Maddala suggest estimating

$$\left(F_t - S_t\right) = \alpha^* + \beta^*\left(S_{t+1} - S_t\right) + w_t. \tag{3}$$

We note however that although the estimate of β^* in equation (3) is consistent asymptotically, it is not unbiased in finite sample if the regressor is endogenous and correlated with w_t. In this respect, one has to take into account that the future spot rate, S_{t+1}, is an endogenous regressor that can be forecasted by its lag,

$$S_{t+1} = \mu + \theta S_t + \eta_{t+1}. \tag{4}$$

The estimation of equation (3) is therefore subject to small sample bias when the error terms, w and η, are correlated. As shown in Lewellen (2004), Mankiw and Shapiro (1986), and Stambaugh (1986, 1999), the finite-sample properties of the OLS estimators can depart substantially from the standard regression setting and can affect inference significantly when the endogenous regressor is correlated with the innovation process. The endogenous regressor, S_{t+1}, in equation (3) may be contemporaneously uncorrelated with the residual, but is not uncorrelated with it at all leads and lags. Following Stambaugh (1999), we find that the small sample bias of the OLS estimate of β^* is

$$E\left(\hat{\beta}^* - \beta^*\right) = \frac{\sigma_{w\eta}}{\sigma_\eta^2} E(\hat{\theta}^* - \theta^*)\ , \tag{5}$$

where $\sigma_{w\eta}$ is the covariance between the error terms. This bias can be large enough to lead to a rejection of the MEH even when it is true. Such a small sample bias can also arise in the standard regression in equation (1) and in studies on the REH, where respectively the forward rate, F_t, and the expected spot exchange rate, S_t^e, are the endogenous regressors.

III. Methodology and Data

For robustness, we use two types of unit root tests to examine the stationarity of the forward rate, the spot exchange rate and its expectation. The first type is the conventional parametric tests, which include the Augmented Dickey-Fuller (ADF) test and the Phillips-Peron (PP) test. We incorporate trends and use the Akaike Information criterion (AIC) to determine the optimal number of lags for these tests. The second type is the Breitung (2002) nonparametric unit root test, which is described in the Appendix. The advantages of the Breitung test are its robustness to structural breaks and alternative model specifications. Hence, there is no need to determine an optimal lag structure. We use the Breitung method to test the null hypothesis of a unit root with drift against the alternative of trend stationary.[2] In addition to these unit root tests, we employ the Engle-Granger (EG) and Johansen tests to examine whether the spot rate is cointegrated with either the expected spot rate or the forward rate.

After determining these cointegrating relationships, we examine the MEH and REH using OLS and Least Absolute Deviations (LAD) regressions. That is, we employ these methods to estimate

$$\log(S_{t+1}) = \alpha + \beta \log(F_t) + u_{t+1} \quad (6)$$

$$\log(S_{t+1}) = \alpha + \beta \log(S_t^e) + \varepsilon_{t+1}, \quad (7)$$

for the MEH and REH, respectively. For the OLS estimation, the Newey-West (1987) and MacKinnon and White (1985) covariance estimators are used to respectively adjust for autocorrelation and heteroscedasticity in the disturbances. The LAD is a median regression method that is preferred to OLS when (i) the data are leptokurtic, (ii) the errors are serially correlated, and (iii) the observations include extreme outliers.[3] It should be noted that none of these regressions are subjected to the well-known spurious regression problem when the series are cointegrated.

[2] We also test the null of a driftless unit root against the alternative of stationarity, but do not report these results to save space.
[3] See Basset and Koenker (1978) for further discussion on the LAD estimation procedure.

To further investigate the MEH, we test for the presence of exchange rate risk premium. As stated in Liu and Maddala (1992), the absence of risk premium implies that the forward rate, F_t, is equal to the expected spot exchange rate, S_t^e, or equivalently

$$\log(F_t) - \log(S_t^e) = \lambda + v_t, \tag{8}$$

where λ is zero under the hypothesis of no risk premium and v_t follows a stationary white noise process. We use the LAD to estimate equation (8) and Breitung (2002) nonparametric unit root test to examine the stationarity of the error term v_t.

As mentioned above, even if all the tests indicate that the series in equations (6) and (7) are stationary and/or cointegrated, the OLS and LAD estimates are still subject to small sample bias. To obtain unbiased OLS and LAD estimates, we use the bootstrapping technique. We determine the bootstrap vector of the response variable for OLS and LAD estimators using the Freedman (1981, 1984) and Hall (1988) resampling method. They suggest resampling of residuals with replacement to solve for the bias in the point estimates when the regressors and the errors are correlated. Due to the random sampling with replacement, the estimates obtained from these resamples vary likewise randomly. As shown in Efron (1987), the sampling distribution of the point estimates will be identical to the population distribution function.[4] The bootstrap estimate of the LAD parameters and the covariance matrix are

$$\hat{\beta}_{LAD} = \frac{\sum_{r=1}^{R} \hat{\beta}_{LAD}(r)}{R} \tag{9}$$

$$Var\left[\hat{\beta}_{LAD}\right] = \frac{1}{R}\sum_{r=1}^{R}\left(\hat{\beta}_{LAD}(r) - \hat{\beta}_{LAD}\right)\left(\hat{\beta}_{LAD}(r) - \hat{\beta}_{LAD}\right)' \tag{10}$$

[4] For a detailed discussion of the bootstrapping technique, see Efron (1979, 1982, and 1987).

where $\hat{\beta}_{LAD}$ is the bootstrap LAD estimator, R is the total number of resamplings with replacements, and $\hat{\beta}_{LAD}(r)$ is the r-th LAD estimate of β. In addition to bias adjusted estimations, we employ bootstrap critical values for the F-tests of the joint hypotheses for the MEH and REH.

As for the data, we use end-of-month observations and forecasts of the exchange rates for the pound sterling, the Deutsche mark, the Japanese yen, the Swiss franc, and the French franc as given by Financial Times Currency Forecaster (FTCF).[5] All exchange rates of these currencies are denominated against U.S. dollar. A currency–forecasting panel, consisting of 30 multinational companies and 15 forecasting services, provides the exchange rate forecasts. The end-of-month forward rates for the same currencies are obtained from the Data Resources Incorporated (DRI) database. The data are monthly and the sample period spans February 1988 to May 2000.

IV. Empirical Results

IV.1. Unit Root and Cointegration Tests

Table 1, Panel A and B, reports the results for the unit root tests for the spot rate, expected spot rate, and forward rates. The ADF and PP tests cannot reject the null of a unit root process for these series. When we repeat these tests for the first-order difference for each series, the null of a unit root process is rejected for all the series. These results are not reported here to save space. In Panel C, the Breitung's (2002) tests indicate for all cases that the unit root hypothesis cannot be rejected. Consistent with most previous

[5] We note hereby that some of these series have been discontinued due to the introduction of the euro as the single currency in the European Monetary Union.

studies, we conclude therefore that the spot rates, expected spot rates and forward rates are $I(1)$ processes.

[Insert Table 1 about here.]

Table 2, Panel A, reports the cointegrating tests and vectors for the spot rate and its expectations. Across all currencies, we find that both the EG and Johansen tests reject the null hypothesis of no cointegration between these the two series. In Panel B, the test results indicate that the spot rate and the forward rate are cointegrated, which concur with the findings in the literature. Using the Johansen method, we also test the hypothesis that the two series are cointegrated with cointegration vector [1, -1]. The LR tests show that the hypothesis is rejected for the pound sterling, the Deutsche mark, and the Japanese yen, while it cannot be rejected for the Swiss franc and the French franc.

[Insert Table 2 about here.]

Since each currency spot rate is cointegrated with its expectations and the forward rate, we conclude that invoking the OLS and LAD regression in (6) and (7) is not subjected to the well-known spurious regression problem.

IV.2. Tests of MEH

Table 3 reports the test results for the efficient market hypothesis. For both the standard OLS and LAD estimation in respectively Panel A and B, we find that the null hypothesis, H_0: $\beta = 1$, is rejected for most of the currencies at the 5% significance level. For all currencies, the OLS and LAD fail to reject the null hypothesis H_0: $\alpha = 0$. Consistent with Hai et al. (1998) and Naka and Whitney (1995), the *F*-tests reject the hypothesis of efficient foreign currency markets, H_0: $\alpha = 0$, $\beta = 1$.

[Insert Table 3 about here.]

In Table 3, Panel C and D, we adjust for the small sample bias and reconstruct the tests using bootstrapped adjusted confidence interval for the hypotheses, H_0: α = 0 and H_0: β = 1. The results indicate a notable difference with the previous results. For both the OLS and LAD estimation, the results show that the value of the bootstrap estimate for β is closer to unity for most currencies when the bootstrapping technique is applied. In addition, we note that for both OLS and LAD the bootstrap estimate of β lies inside the percentile bootstrap confidence interval for most currencies. For the LAD estimation, the bootstrap estimate of β can be considered as insignificantly different from unity for all the currencies, except the Deutsche mark.

The bias adjustment also has an impact on the inference about market efficiency. As shown in Panel C and D, the critical values of the bootstrap *F* distribution are shifted to the right relative to the asymptotic *F* distribution. For the LAD estimation, the *F*-tests based on the bootstrap critical values indicate that the joint hypothesis of market efficiency, H_0: $\alpha = 0$, $\beta = 1$, cannot be rejected for the Swiss franc and French franc. For the OLS estimation, we find that the MEH can no longer be rejected for the French franc. For the other currencies, the question remains whether the MEH is rejected because of a failure of the REH, the presence of a risk premium, or both.

IV.3. Tests of REH

Table 4, Panel A and B, presents the results for the REH using the standard OLS and LAD estimation methods. We observe that most of the LAD estimates for β are closer to unity as compared to the OLS estimates. For both the OLS and LAD estimation, the *t*-tests reject the null hypothesis for the slope coefficient, H_0: $\beta = 1$, for the pound sterling, Deutsche mark, Swiss franc, and French franc. The *F*-tests based on both estimation methods indicate that the joint hypothesis of rational expectations, H_0: $\alpha = 0$, $\beta = 1$, is rejected for the pound sterling and Swiss franc.

[Insert Table 4 about here.]

Panel C and D report the bias adjusted OLS and LAD estimates using the bootstrap method. We notice that the role of small sample bias is even more significant for the REH. For all currencies, the OLS and LAD bootstrap estimate of β is closer to unity than the standard estimates. Using the bootstrapped adjusted confidence interval in Panel C and D, we find that across all currencies, we cannot reject the hypotheses for the point estimates, H_0: $\alpha = 0$ and H_0: $\beta = 1$. The impact of the bias adjustment become more evident when we test the joint hypothesis, H_0: $\alpha = 0$, $\beta = 1$. Based on the bootstrap critical values, the *F*-tests indicate that across all currencies the REH cannot be rejected after the bias adjustment. In contrast to Liu and Maddala (1992), we conclude that the REH cannot be rejected for the one-month horizon.

IV.4. The risk premium

Table 5 reports the results of the tests of the risk premium hypothesis. Consistent with Liu and Maddala (1992), Naka and Whitney (1995), and Wu and Chen (1998), we find evidence supporting the presence of risk premium for most currencies. The parameter estimate for λ is statistically significant for the pound sterling, Deutsche mark, and Japanese yen. As mentioned above, the hypothesis of no risk premium requires v_t in equation (8) to follow a stationary white noise process or, equivalently, the forward rate and the expected spot rate to be cointegrated. For the Pound sterling, the Deutsche mark, and the Japanese yen, we observe that the Breitung tests cannot reject the unit root hypothesis for the innovation process in equation (8), providing thereby evidence in support of a foreign exchange rate risk premium for these currencies.

On the other hand, the Breitung tests cannot reject the hypothesis of no risk premium for the Swiss franc and French franc. The absence of both risk premium and expectation errors explains why the MEH is accepted for these two currencies after the biased adjustment. For the other currencies, we infer that the market efficiency hypothesis is rejected because of the existence of a foreign exchange risk premium.

V. Conclusion

We have examined the role of small sample bias in the inference about the foreign exchange market efficiency hypothesis for the pound sterling, Deutsche mark, Japanese yen, Swiss franc, and French franc. We have employed the Least Absolute Deviation estimation to account for the fat tails of exchange rates distribution and bootstrapping technique to adjust for the small sample bias. When we use standard methods, we find that the efficiency hypothesis is rejected for all five currencies under consideration. However, after adjusting for small sample bias, we find that the hypothesis cannot be rejected for the Swiss franc and French franc. This result suggests that the bias is large enough to affect inference and that standard methods reject the efficiency hypothesis too often.

We also investigate the impact of small sample bias on the rational expectation hypothesis, which is one of the explanations for the failure of market efficiency. We find that the role of small sample bias is even more significant for the rational expectation hypothesis. After bias adjustment, our results indicate that rationality cannot be rejected for the monthly data. In addition, we find evidence supporting the existence of foreign exchange rate risk premium for the pound sterling, Deutsche mark, and Japanese yen. We infer therefore that the failure of market efficiency hypothesis for these currencies is not related expectation errors, but rather to the presence of the risk premium.

Appendix

Consider a time series y_t that is regressed on a vector z_t, where $\hat{u}_t$ are the regression residuals. The typical elements of z_t are a constant, a time trend or dummy variables. The Breitung's (2002) nonparametric method to test the null hypothesis that y_t is $I(1)$ against the alternative $y_t \sim I(0)$ is based on the following variance ratio statistic:

$$Q_T = \frac{T^{-1}\sum_{t=1}^{T}\hat{U}_t^2}{\sum_{t=1}^{T}\hat{u}_t^2}, \tag{11}$$

where $\hat{U}_t = \hat{u}_1 + ... + \hat{u}_t$ denotes the partial sum process. The variance ratio statistic is a left tailed test that rejects the null hypothesis when the test statistic falls below the corresponding critical value. Under the null hypothesis, the test statistic, $T^{-1}Q_T$, converges in distribution to a function of a standard Brownian motions,

$$T^{-1}Q_T \Rightarrow \frac{\int_0^1 \left[\int_0^a \tilde{W}_j(s)ds\right]^2 da}{\int_0^1 \tilde{W}_j(a)^2 da}, \tag{12}$$

where

$$\tilde{W}_0(s) \equiv W(s) \quad \text{for } z_t = 0,$$

$$\tilde{W}_1(s) \equiv W(s) - \int_0^1 W(a)^2 da \quad \text{for } z_t = 1,$$

$$\tilde{W}_3(s) \equiv W(s) - (4-6s)\int_0^1 W(a)da - (12s-6)\int_0^1 aW(a)da \quad \text{for } z_t = [1, t]'.$$

It should be noted that this asymptotic distribution is free of nuisance parameters. Under the stationary and trend stationary alternative hypotheses, we have that $T^{-1}Q_T \Rightarrow 0$ as T goes to infinity.

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Table 1

Unit root tests with a linear trend in the fitted regressions

	Currencies				
Tests	£	DM	¥	SW	FF
Panel A: Test results for the logarithm of spot exchange rates					
ADF	-2.789	-1.894	-1.245	-2.251	-2.109
	(0.200)	(0.657)	(0.901)	(0.461)	(0.540)
PP	-15.978	-9.081	-6.098	-12.970	-11.115
	(0.154)	(0.500)	(0.737)	(0.267)	(0.365)
Breitung	0.007	0.010	0.014	0.007	0.009
Panel B: Test results for the logarithm of expected spot exchange rates					
ADF	-2.841	-2.178	-0.171	-2.643	-2.423
	(0.182)	(0.502)	(0.916)	(0.260)	(0.367)
PP	-16.357	-10.438	-6.131	-15.027	-13.362
	(0.143)	(0.407)	(0.735)	(0.184)	(0.250)
Breitung	0.008	0.01	0.014	0.006	0.009
Panel C: Test results for the logarithm of forward exchange rates					
ADF	-1.696	-1.548	-1.234	-2.093	-1.980
	(0.752)	(0.793)	(0.903)	(0.550)	(0.611)
PP	-6.680	-6.015	-5.082	-9.535	-9.281
	(0.691)	(0.780)	(0.815)	(0.468)	(0.486)
Breitung	0.007	0.017	0.023	0.007	0.006

Notes: This table reports the test statistics for the ADF, PP, and Breitung unit root tests applied to the logarithm of the spot exchange rate, expected spot exchange rate, and forward exchange rate. The data are monthly and the sample period spans from February 1988 to May 2000. The *p*-values for the ADF and PP tests are reported in parentheses. These asymptotic *p*-values are computed using MacKinnon (1994) approximation, which is robust to finite sample distortion. Based on the Akaike Information criterion, the optimal number of lags for the ADF and PP tests is equal to 2 for all currencies, except being equal to 6 for the Japanese yen (¥). The Breitung statistics test the null hypothesis of a unit root with drift against the alternative of trend stationary. The Breitung's (2002) test is described in the Appendix. The null hypothesis of unit root is rejected at the 5% significance level when the test statistic falls below the critical value of 0.00342. Consistent with Breitung (2002), we use Gaussian random walk sequences to compute this critical value from the empirical distribution of 10,000 realizations of the limiting expression of the test statistic in the Appendix.

Table 2

Engle-Granger (EG) and Johansen (J) cointegration tests

	Currencies				
	£	DM	¥	SW	FF
Panel A: The logarithm of future spot exchange rate and its expectations					
EG	-6.502	-6.084	-0.388	-5.889	-4.756
	(0.000)	(0.000)	(0.003)	(0.002)	(0.021)
Number of lags	2	2	7	2	5
Cointegrating vector	[1.0, -0.874]	[1.0, -0.944]	[1.0, -0.977]	[1.0, -0.902]	[1.0, -0.926]
J-trace statistic, $r = 0$	74.7*	61.9*	71.1*	81.0*	94.8*
J-trace statistic, $r = 1$	7.1	4.0	2.5	5.6	4.7
Cointegrating vector	[1, -0.9803]	[1, -0.9765]	[1, -0.9705]	[1, -0.9744]	[1, -0.9837]
Panel B: The logarithm of future spot exchange rate and forward exchange rate					
EG	-6.501	-5.819	-5.472	-5.859	-5.867
	(0.000)	(0.002)	(0.0001)	(0.000)	(0.000)
Number of lags	2	3	3	3	2
Cointegrating vector	[1.0, -0.809]	[1.0, -0.921]	[1.0, -0.955]	[1.0, -0.884]	[1.0, -0.954]
J-trace statistic, $r = 0$	904.3*	65.4*	652.3*	779.4*	90.1*
J-trace statistic, $r = 1$	8.8	4.0	2.9	3.7	4.6
Cointegrating vector	[1, -0.9851]	[1, -0.9802]	[1, -0.9940]	[1, -0.9986]	[1, -0.9807]

Notes: This table presents the Engle-Granger (EG) tests for whether the logarithm of the future spot exchange rate is cointegrated with the logarithm of its expectations and the logarithm of the forward rate. We use the Akaike Information criterion to determine the optimal number of lags for these tests. The *p*-values and cointegrating vectors are reported respectively in parentheses and square brackets. Since the null hypothesis for the EG tests is no cointegration, these *p*-values imply a rejection of no integration for all the cases. For the Johansen (J) tests, we impose cointegrating restrictions on the intercept parameters. The J-trace statistic is the likelihood ratio for the Johansen cointegration test. This test starts with the null hypothesis $r = 0$, where r represents the cointegrating rank. The asterisk * indicates rejection of the null hypothesis of no cointegration at 5% significance level.

Table 3

Parameter estimates and tests of the efficient market hypothesis:

$$\log(S_{t+1}) = \alpha + \beta \log(F_t) + u_{t+1}$$

	Currencies				
	£	DM	¥	SW	FF
Panel A: Standard OLS estimates					
α	0.031	0.041	-0.194	-0.022	-0.084
	(3.897)*	(2.148)*	(-3.592)*	(-4.297)*	(-1.555)
β	0.934	0.921	0.960	0.953	0.984
	(-3.642)*	(-2.088)*	(-3.541)*	(-4.103)*	(-1.545)
F-statistic	10.081*	6.732*	8.557*	9.259*	4.627*
Panel B: Standard Least Absolute Deviations (LAD) estimates					
α	0.025	0.043	-0.209	-0.013	-.0167
	(8.673)*	(3.080)*	(-9.773)*	(-1.587)	(-2.077)*
β	0.947	0.913	0.957	0.944	0.992
	(-8.760)*	(-3.120)*	(-9.548)*	(-2.327)*	(-1.651)
F-statistic	18.392*	11.963*	17.168*	11.287*	10.560*
Panel C: Bias adjusted OLS method					
Boot α	0.027	0.032	-0.219	-0.019	0.028
	[-0.008, 0.010]	[-0.037, 0.034]	[-0.057, 0.060]	[-0.019, 0.023]	[-0.019, 0.020]
Boot β	0.998	0.923	0.954	0.965	0.984
	[0.922, 0.963]	[0.976, 0.995]	[0.942, 0.967]	[0.953, 0.977]	[0.973, 0.996]
Boot *F*	4.446	2.877	3.420	4.561	5.058†
Panel D: Bias adjusted LAD method					
Boot α	0.001	0.049	-0.218	0.024	0.002
	[-0.008, 0.011]	[-0.037, 0.047]	[-0.049, 0.047]	[-0.007, 0.005]	[-0.024, 0.037]
Boot β	0.961	0.904	0.954	0.985	1.001
	[0.938, 0.982]	[0.906, 1.074]	[0.944, 0.964]	[0.949, 0.985]	[0.978, 1.014]
Boot *F*	5.051	8.241	4.650	15.598†	12.460†

Notes: Panel A and B respectively provide standard OLS and LAD estimates for the market efficiency hypothesis. For the OLS estimation, the Newey-West (1987) and MacKinnon and White (1985) covariance estimators are used to respectively adjust for autocorrelation and heteroscedasticity in the disturbances. The LAD is a median regression method that provides robust estimators in the presence of leptokurtic distributions and extreme outliers. The numbers in the parentheses are the *t*-test statistics for H_0: $\alpha = 0$ and H_0: $\beta = 1$. The *F*-statistics test the market efficiency hypothesis, H_0: $\alpha = 0$, $\beta = 1$. The critical value for the *F*-tests is equal is 3.064. An asterisk * indicates that the null hypothesis is rejected at 5% significance level. Panel C and D presents bootstrap OLS and LAD estimates for the market efficiency hypothesis. The numbers in the square brackets are the bootstrap confidence intervals. The upper tail of the interval is determined at the conventional 2.5% level and the lower tail is at 97.5%. The boot *F* is the bootstrap critical value for the *F*-tests of the joint hypothesis, H_0: $\alpha = 0$, $\beta = 1$. For the *F*-tests, † indicates that the point estimates are jointly inside the bootstrap interval and the null hypothesis cannot be rejected at 5% significance level.

Table 4

Parameter estimates and tests of the rational expectation hypothesis:

$$\log(S_{t+1}) = \alpha + \beta \log(S_t^e) + \varepsilon_{t+1}$$

	Currencies				
	£	DM	¥	SW	FF
Panel A: Standard OLS estimates					
α	-0.061	0.024	-0.087	0. 035	0.187
	(-2.578)*	(1.343)	(0.740)	(3.052)*	(2.244)*
β	0.878	0.950	0 .981	0.900	0. 891
	(-2.377)*	(-1.405)	(-0.745)	(-3.046)*	(-2.233)*
F-statistic	4.810*	1.025	0.292	3.782*	2.550
Panel B: Standard Least Absolute Deviations (LAD) estimates					
α	-0.060	0.019	-0.034	0.048	0.161
	(-4.418)*	(1.499)	(-0.588)	(5.039)*	(3.291)*
β	0.883	0.962	1.010	0.869	0.910
	(-4.317)*	(-1.520)	(0.601)*	(-5.007)	(-3.299)*
F-statistic	6.817*	1.155	0.293	7.781*	2.997
Panel C: Bias adjusted OLS method					
Boot α	0.000	0.000	0.000	0.000	0.001
	[-0.038, 0.036]	[-0.033, 0.032]	[-0.240, 0.218]	[-0.026, 0.024]	[-0.130, 0.130]
Boot β	1.000	1.001	1.000	0.999	1.002
	[0.923, 1.072]	[0.934, 1.061]	[0.953, 1.049]	[0.930, 1.071]	[0.923, 1.074]
Boot *F*	5.015†	3.386†	3.608†	3.951†	4.306†
Panel D: Bias adjusted LAD method					
Boot α	-0.028	0.023	-0.0309	0.041	0.014
	[-0.035, 0.038]	[-0.027, 1.014]	[-0.242, 0.318]	[-0.040, 0.130]	[-0.156, 0.116]
Boot β	1.002	0.947	1.007	1.017	1.012
	[1.001, 1.883]	[0.944, 0.988]	[1.000, 1.050]	[0.996,1.024]	[0.906, 1.090]
Boot *F*	7.190†	8.494†	8.958†	10.078†	7.613†

Notes: Panel A and B respectively provide standard OLS and LAD estimates for the rational expectations hypothesis. For the OLS estimation, the Newey-West (1987) and MacKinnon and White (1985) covariance estimators are used to respectively adjust for autocorrelation and heteroscedasticity in the disturbances. The LAD is a median regression method that provides robust estimators in the presence of leptokurtic distributions and extreme outliers. The numbers in the parentheses are the *t*-test statistics for H_0: $\alpha = 0$ and H_0: $\beta = 1$. The *F*-statistics test the market efficiency hypothesis, H_0: $\alpha = 0$, $\beta = 1$. The critical value for the *F*-tests is equal is 3.064. An asterisk * indicates that the null hypothesis is rejected at 5% significance level. Panel C and D presents bootstrap OLS and LAD estimates for the rational expectation hypothesis. The numbers in the square brackets are the bootstrap confidence intervals. The upper tail of the interval is determined at the conventional 2.5% level and the lower tail is at 97.5%. The boot *F* is the bootstrap critical value for the *F*-tests of the joint hypothesis, H_0: $\alpha = 0$, $\beta = 1$. For the *F*-tests, † indicates that the point estimates are jointly inside the bootstrap interval and the null hypothesis cannot be rejected at 5% significance level.

Table 5

Stationarity tests for the presence of risk premium:

$$\log(F_t) - \log(S_t^e) = \lambda + v_t$$

	Currencies				
	£	DM	¥	SW	FF
λ	0.001 (0.001)	0.001 (0.000)	0.007 (0.000)	0.001 (0.653)	0.001 (0.118)
Breitung	0.01302	0.01122	0.02611	0.00112*	0.00176*

Notes: This table reports the LAD estimates for the parameter λ. The numbers in the parentheses are the *p*-values for the null hypothesis of zero risk premium, H_0: $\lambda = 0$. The Breitung statistics test the null of a driftless unit root against the alternative of stationarity. The null hypothesis of unit root is rejected at the 5% significance level when the test statistic falls below the critical value of 0.01004. Consistent with Breitung (2002), we use Gaussian random walk sequences to compute this critical value from the empirical distribution of 10,000 realizations of the limiting expression of the test statistic in the Appendix. An asterisk * indicates that the null hypothesis cannot be rejected at 5% significance level.