DOCUMENTO DE TRABAJO

9130

FURTHER TESTS ON THE FORWARD EXCHANGE RATE UNBIASEDNESS HYPOTHESIS

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I.-INTRODUCTION.

There is a large and growing literature devoted to the question of whether the forward market for foreign exchange is efficient.

The original concept of an efficient market is due to Fama who defined such a market as "a market where there are large numbers of rational, profit-maximizers actively competing, with each trying to predict future market values of individual securities, and where important current information is almost freely available to all participants" (Fama, 1965, p. 56). Thus, in an efficient market security prices at any time should fully reflect all available information and no profit opportunities are left unexploited.

In the foreign exchange market, if participants are rational and risk neutral, expectations concerning future rates should be incorporated and reflected in forward exchange rates. Thus the forward exchange rate should be an unbiased and efficient predictor of the future spot rate. Hence a regression of the observed spot rate at time t+k (s_{t+k}) on the forward rate determined at time t for settlement k-periods ahead (f_t^k) (where exchange rates are measured by natural logarithms of currency prices of foreign exchange),

$$s_{t+k}=\alpha+\beta f_t^k+\varepsilon_{t+k}, \qquad (1)$$

should result in an estimated constant ($\hat{\alpha}$) not significantly different from zero and an estimated coefficient on the forward rate ($\hat{\beta}$) not significantly different from one.

This proposition is known as the <u>forward rate unbiasedness hypothesis</u>. The forward rates in this formulation are regarded as the directly observable expectations of the spot exchange rates (see Frenkel, 1977, 1980).

For a number of currencies and time periods the overwhelming findings is that the forward exchange rate is an inefficient predictor of the spot rate (see Hodrick, 1987, and Baillie and McMahon, 1989 for a further discussion). This conclusion is based on traditional estimation and inference procedures, which are only applicable if the variables in the regression model (1) are stationary. However, the increasing evidence that the spot and the forward rates are not stationary variables (see, e. g. Meese and Singleton, 1982; and Baillie and Bollerslev, 1987) has led to the need of reexamining such conclusion. To properly account for the non-stationarity of the series, Burridge and Ngama (1990) have recently recommended to use the general asymptotic framework for estimation and inference in regressions with integrated variables developed by Phillips and Hansen (1990). In this paper we apply this framework to reexamine the evidence on the forward rate unbiasedness hypothesis for the main currencies vis-à-vis the U. S. Dollar exchange rate using weekly data.

We use Barclays Bank's quotations for weekly spot, one-, two-, three- and six-months forward exchange rates collected by

Datastream Company. The data cover the period 1st week November 1983 to 1st week December 1987, giving a total of 214 observations on the Pound Sterling/U. S. Dollar, Deutschemark/U. S. Dollar, French Franc/U. S. Dollar and Swiss Franc/U. S. Dollar exchange rates. For the Japanese Yen/U. S. Dollar exchange rates, the data cover the period 1st week of April 1984 to 4th week October 1987.

The paper is organized as follows. In Section II we test for unit root and cointegration. Section III describes the salient features of the Phillips and Hansen procedure. We discuss our results in Section IV. Concluding remarks are given in Section V.

II.-INTEGRATION AND COINTEGRATION TESTS.

Before applying the Phillips and Hansen (1990) procedure it is necessary to determine the order of integration of the variables, and to test whether the future spot exchange rates are cointegrated with their respective forward rates.

A) UNIVARIATE TIME SERIES PROPERTIES.

Several statistical tests for unit roots have been developed to test for stationarity in time series[1]. In this study we use the Phillips and Perron (1988) robust tests.

Phillips and Perron consider three alternative data generating models:

$$Y_t=\tilde{\mu}+\tilde{\beta}(t-T/2)+\tilde{\alpha}Y_{t-1}+\tilde{\varepsilon}_t \qquad (2)$$

$$Y_t=\mu^*+\alpha^*Y_{t-1}+\varepsilon^*_t \qquad (3)$$

and

$$Y_t=\hat{\alpha}Y_{t-1}+\hat{\varepsilon}_t \qquad (4)$$

where T is the sample size.

Their approach to testing for unit roots involves estimating the coefficients of these equations by ordinary least squares and constructing the non-parametrically corrected test statistics $Z(t_{\tilde{\alpha}})$, $Z(t_{\tilde{\beta}})$, $Z(\phi_3)$, and $Z(\phi_2)$ to test the null hypotheses H^1_0: $\tilde{\alpha}=1$, H^2_0: $\tilde{\beta}=0$, H^3_0:$(\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (\tilde{\mu}, 0, 1)$, and H^4_0: $(\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (0, 0, 1)$ in equation (2) respectively; the statistics $Z(t_{\alpha^*})$, $Z(t_{\mu^*})$ and $Z(\phi_1)$

to test the null hypotheses H_0^5: $\alpha^*=1$, H_0^6: $\mu^*=0$ and H_0^7:$(\mu^*, \alpha^*) = (0, 1)$ in equation (3) respectively; and the statistic $Z(t_{\hat{\alpha}})$ to test the null hypothesis H_0^8: $\hat{\alpha}=1$ in equation (4)[2].

In making inferences we follow the testing sequence suggested in Perron (1988), with the modifications presented in Dolado *et al.* (1990) and Sosvilla-Rivero (1990). This testing sequence is based on the idea that the most plausible alternatives to the unit root hypothesis are stationarity, or stationarity about a linear trend, or either of these with drift. Because the null distributions of unit root tests are not invariant to the presence of trend, we start the sequence with the most general model containing a drift and a trend as a regressors, and we use the test statistics $Z(\phi_3)$ and $Z(t_{\tilde{\alpha}})$ to assess whether there is evidence for rejecting the null hypothesis of a unit root in a regression with a fitted drift *and* trend. If it is rejected there is no need to go further.

If the null hypothesis is not rejected, this may be due to the poor power properties of these statistics compared to those from a regression containing a constant alone, allowing for a non-zero mean in the series. To check the validity of the latter, we use $Z(t_{\tilde{\beta}})$ test for the significance of the trend under the null[3]. If it is significant, the we test again for a unit root using the standardized normal distribution[4]. If the trend is *not* significant in the maintained model, we test again for the presence of a unit root with a drift using the test statistics $Z(\phi_2)$ and $Z(t_{\alpha^*})$. If this null is rejected, again there is no need to go further.

Finally, if this null is not rejected, we test for the significance of the drift under the null of a unit root using the test statistic $Z(t_{\mu}*)$. If the drift is significant we test again for a unit root using the standardized normal. If the drift is not significant in the maintained model, then to maximize the power of our unit root test, we use the test statistics $Z(\phi_1)$ and $Z(t_{\hat{\alpha}})$.

Following Dickey and Pantula (1987), we start testing for integration of order two [the null is I(3)] on the second difference of the variables and go down testing for integration of order one [the null is I(2)] on the first difference of the variables and for integration of order zero [the null is I(1)] on the level of the variables.

Using Peter Burridge's ROOTINE programme, with a four period lag in the Newey-West variance estimator, we computed the Phillips-Perron tests for all the series. The results for second and first differences of the series showed that the null hypotheses of I(3) and I(2) processes were rejected at the 1% level of significance[5]. Table 1 presents the results for the levels of all the exchange rates considered. We find that the null of a single unit root cannot be rejected for any of the spot or forward exchange rate under study.

B) COINTEGRATION TESTS.

Once that we have found that all the variables are integrated of order one [I(1)], we test for cointegration between the future spot exchange rates and their respective forward rates by testing the residuals from equation (1) for unit roots. Phillips and Ouliaris (1990) have recently proposed a modification of the augmented Dickey-Fuller statistic for the cointegrating residuals (CRADF statistics) based on Phillips and Perron tests. We denote by $\hat{Z}_t$ std, $\hat{Z}_t$ μ and $\hat{Z}_t$ $\mu\beta$ their residual-based test for the null hypothesis of no cointegration in the standard, demeaned, and demeaned and detrended cases. Critical values for these tests are provided in Phillips and Ouliaris (1990, p. 192).

Tables 2 reports the cointegration tests[6]. Only for the one- and two-month forward maturity are the cointegration tests significant for all currencies and, therefore, we are able to reject the null of non-cointegration. In the case of the three-moth maturity, only the French Franc's cointegration tests for the standard regression are significant.

III.- STATISTICAL INFERENCE IN REGRESSION WITH I(1) VARIABLES.

Note that the joint dependence of most aggregate time series and their non-stationarity invalidate the routine application of many standard statistical procedures in the cointegrating regression. Phillips and Hansen (1990) present a class of Wald tests which are modified by semiparametric corrections for serial correlation and for endogeneity. The resulting test statistics (termed fully-modified Wald tests) have limiting χ^2 distributions and therefore allow inference to proceed in a conventional way.

Basically the application of the Phillips and Hansen procedure involves the following[7]:

A) CONSTRUCTION OF THE LONG RUN COVARIANCE MATRIX.

The main impediment to inference in a cointegrating regression is that the covariance matrix as conventionally calculated is not consistent. A consistent estimate is given by the "long run" covariance matrix, which is the same as the Phillips and Durlauf (1986) heteroskedasticity and serial correlation consistent covariance matrix. To construct this covariance matrix we need estimates of the innovation vector $\xi_t = (\varepsilon_{t+k}, \nu_t)$, where $\nu_t = \partial f^k_t$. Phillips and Durlauf (1986) showed that the asymptotic distributions of $\hat{\alpha}$ and $\hat{\beta}$ depend on the long run covariance matrix of ξ_t given as[8]:

$$\Omega = \Sigma_0 + \sum_{j=1}^{k}\left[\Lambda_j + \Lambda'_j \right] , \qquad (5)$$

where $\Sigma_0 = \lim_{t\to\infty} T^{-1}\sum_{t=1}^{T}E(\xi_t\xi'_t)$, $\Lambda_j = \lim_{t\to\infty} T^{-1}\sum_{t=j+1}^{T}E(\xi_t\xi'_{t-j})$, T is, as before, the total number of observations, and k is the order of serial correlation to be corrected for.

The estimate of the long run covariance matrix will also be used in making the necessary adjustments that will allow bias-corrected estimation of the cointegrating vector and the construction of fully modified test statistics. To do this the long run covariance matrix has to be partitioned into its various sub-matrices as follows:

$$\Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix}. \tag{6}$$

Another matrix which will be needed for the asymptotic bias correction is

$$\Delta = \Sigma_0 + \Lambda_J. \tag{7}$$

The matrix Δ can also be partitioned into its various sub-matrices as in (6). Note that, for $\hat{\Omega}$ to be positive, semi-definite covariance smoothing techniques, commonly known as Newey and West (1987) adjustment, must be used in constructing it.

B) LONG RUN ENDOGENEITY CORRECTION.

Phillips and Hansen found that the limit distributions of the usual Wald statistics for testing hypotheses on the cointegrating vector are non-standard and depend on nuisance parameters. These nuisance parameter dependencies arise because the long run coefficient of determination between $\{\varepsilon_{t+k}\}$ and $\{\nu_t\}$ is non-zero. They therefore suggested that the dependent variable be adjusted so that the resulting disturbance term will be long run uncorrelated with ν_t. In our case the adjusted dependent variable will be:

$$s^{+}_{t+k} = s_{t+k} - \hat{\Omega}_{12}\hat{\Omega}^{-1}_{22}\partial f^{k}_{t} \qquad (8)$$

This correction will result in the long-run covariance matrix for the cointegrating regression being

$$\begin{bmatrix} \Omega_{11.2} & 0 \\ 0 & \Omega_{22} \end{bmatrix}, \qquad (9)$$

where $\hat{\Omega}_{11.2} = \hat{\Omega}_{11} - \hat{\Omega}_{12}\hat{\Omega}^{-1}_{22}\hat{\Omega}_{21}$.

C) ESTIMATION AND TESTING.

Estimating equation (1) now involves regressing s^{+}_{t+k} on a constant and f^{k}_{t}. However, the estimate of the cointegrating vector, $\hat{\Gamma}^{+} = [\hat{\alpha}^{+}, \hat{\beta}^{+}]$, is still asymptotically biased. Phillips and Hansen suggested the modified (bias-corrected) estimator

$$\hat{\Gamma}^{+*} = \hat{\Gamma}^{+} - \hat{T}J\hat{\Delta}^{+}_{21}(X'X)^{-1} \qquad (10)$$

where $J' = \begin{array}{c} 1 \\ \begin{bmatrix} 0 \\ I \end{bmatrix}\begin{matrix} 2 \\ 1 \end{matrix} \end{array}$, $\hat{\Delta}^{+}_{21} = [\hat{\Delta}_{21} - \hat{\Omega}_{12}\hat{\Omega}\hat{\Delta}_{22}]$ and $X = [1, f^{k}_{t}]$.

Finally, they showed that

$$T^{-1}\left(\hat{\Gamma}^{+*} - \Gamma\right) \sim N\left(0, [\Omega_{11.2} \otimes (X'X)^{-1}]\right). \qquad (11)$$

Now, a <u>fully modified Wald statistic</u> can be constructed to test the null hypothesis:

$$H_0: R \text{ vec } \Gamma = r,$$

where, when testing whether an individual parameter is zero, r=0 and $R=I_p$. In particular, under this null hypothesis the Wald statistic:

$$W(1) = (R \text{ vec } \hat{\Gamma}^{+*} - r)' \; [R((\hat{\Omega}_{11.2} \otimes (X'X)^{-1})R'] \; (R \text{ vec } \hat{\Gamma}^{+*} - r) \sim \chi^2_q$$

where the degrees of freedom q is the number of restrictions (in this case q=1).

IV.-Empirical Results.

Tables 3 and 4 present the results of estimating equation (1) using the Phillips and Hansen fully modified procedure[9]. Given the results from our cointegration tests, we only report the results for one- and two-month maturities.

The results for one-month forward maturity show that the forward rate unbiasedness hypothesis (FUH) cannot be rejected for all the currencies we considered in this study. For the two-month forward maturity, the fully modified Wald statistics for testing $\alpha=0$ [$W_{\alpha}(1)$] for the U. K. and Germany are significant. Thus we reject FUH for those currencies, while we are unable to reject it for Switzerland, France and Japan.

V.- Concluding Remarks

The foreign exchange market is said to be "efficient" if the exchange rates fully and instantly reflect all available information and no profit opportunities are left unexploited. Inherent in this relationship are the assumptions of rational expectations, risk neutrality and costless transactions.

Most of the attempts to evaluate empirically the efficiency of the forward exchange markets have concentrated on whether the forward exchange rate is an unbiased and efficient predictor of the future spot rate using equation (1).

Given the presence of unit roots in the series, we applied the estimation and inference procedures developed for I(1) processes by Phillips and Hansen (1990) for testing the forward rate unbiasedness hypothesis using weakly data for the main exchange rates. Our results are consistent with that hypothesis for one-month forward maturity. In the case of two-month forward maturity, only the results for the U. K. Pound/U. S. Dollar and the Deutschemark/U. S. Dollar reveal that the forward rate is not an unbiased predictor of the future spot rate. For other maturities, the requirement of equation (1) to be a cointegrating regression was not met.

ACKNOWLEDGEMENTS

The authors are extremely grateful to Peter Burridge for the programme used in this paper and for helpful comments.

FOOTNOTES:

[1]. See Dolado *et al.* (1990) for a survey.

[2]. See Perron (1988) for a definition of these tests.

[3]. Note, however, that, as Perron (1988) points out, $Z(t_{\tilde{\beta}})$ is not invariant with respect to m, so can only be used if $Z(\phi_2)$ suggests the absence of a drift.

[4]. See West (1988) for a discussion of the use of the standardized normal distribution tables instead of the Fuller (1976) tables when the drift or trend is significant under the null.

[5]. The results of the three and two unit roots are not reported here for reasons of space. They can, however, be obtained from the authors.

[6]. The number of lags in the Newey-West variance estimator is 4, 8, 13 and 26 for one-, two-, three- and six-month forward maturity, respectively.

[7]. See Phillips and Hansen (1990) for a detailed explanation of this procedure.

[8]. To obtain a positive semi-definite estimate of $\hat{\Omega}$, we constructed it using the covariance smoothing technique suggested by Newey and West (1987).

[9]. The number of lags in the Newey-West variance estimator is 4 and 8 for one- and two-month forward maturity, respectively.

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TABLE 1: PHILLIPS-PERRON TESTS FOR UNIT ROOTS.

A) U. K. Pound.

	Spot	1-MFwd	2-MFwd	3-MFwd	6-MFwd
$Z(\Phi_3)$	3.76	3.72	3.69	3.64	3.54
$Z(t_{\tilde{\alpha}})$	-1.90	-1.88	-1.85	-1.83	-1.76
$Z(t_{\tilde{\beta}})$	2.76	2.74	2.73	2.71	2.67
$Z(\Phi_2)$	2.71	2.68	2.66	2.62	2.53
$Z(t_{\alpha^*})$	-0.18	-0.19	-0.21	-0.24	-0.27
$Z(t_{\mu^*})$	0.40	0.41	0.42	0.44	0.47
$Z(\Phi_1)$	0.30	0.29	0.28	0.27	0.25
$Z(t_{\hat{\alpha}})$	0.69	0.67	0.65	0.63	0.60

B) Deutschemark.

	Spot	1-MFwd	2-MFwd	3-MFwd	6-MFwd
$Z(\Phi_3)$	3.97	3.94	3.93	3.92	3.85
$Z(t_{\tilde{\alpha}})$	-2.01	-2.01	-2.00	-1.98	-1.95
$Z(t_{\tilde{\beta}})$	-2.75	-2.74	-2.74	-2.73	-2.70
$Z(\Phi_2)$	3.84	3.79	3.79	3.77	3.71
$Z(t_{\alpha^*})$	0.67	0.65	0.66	0.68	0.67
$Z(t_{\mu^*})$	-1.06	-1.05	-1.06	-1.06	-1.06
$Z(\Phi_1)$	2.09	1.99	2.00	2.00	1.99
$Z(t_{\hat{\alpha}})$	-1.61	-1.59	-1.59	-1.59	-1.58

C) French Franc

	Spot	1-MFwd	2-MFwd	3-MFwd	6-MFwd
$Z(\Phi_3)$	3.83	3.87	3.91	3.93	3.84
$Z(t_{\tilde{\alpha}})$	-2.14	-2.14	-2.14	-2.14	-2.12
$Z(t_{\tilde{\beta}})$	-2.76	-2.77	-2.78	-2.79	-2.75
$Z(\Phi_2)$	3.19	3.35	3.40	3.43	3.43
$Z(t_{\alpha^*})$	0.34	0.36	0.39	0.40	0.43
$Z(t_{\mu^*})$	-0.46	-0.49	-0.51	-0.53	-0.56
$Z(\Phi_1)$	1.25	1.24	1.27	1.31	1.40
$Z(t_{\hat{\alpha}})$	-1.42	-1.44	-1.46	-1.47	-1.52

TABLE 1 (concluded).

D) Swiss Franc.

	Spot	1-MFwd	2-MFwd	3-MFwd	6-MFwd
$Z(\Phi_3)$	4.02	3.98	3.94	3.91	3.87
$Z(t_{\tilde{\alpha}})$	-1.97	-1.96	-1.94	-1.93	-1.87
$Z(t_{\tilde{\beta}})$	-2.77	-2.76	-2.74	-2.73	-2.71
$Z(\Phi_2)$	3.69	3.64	3.61	3.60	3.53
$Z(t_{\alpha^*})$	0.64	0.62	0.62	0.61	0.63
$Z(t_{\mu^*})$	-1.09	-1.07	-1.07	-1.06	-1.07
$Z(\Phi_1)$	1.80	1.74	1.74	1.71	1.71
$Z(t_{\hat{\alpha}})$	-1.42	-1.40	-1.39	-1.38	-1.37

E) Japanese Yen.

	Spot	1-MFwd	2-MFwd	3-MFwd	6-MFwd
$Z(\Phi_3)$	3.79	3.71	3.72	3.73	3.71
$Z(t_{\tilde{\alpha}})$	-2.43	-2.40	-2.40	-2.39	-2.37
$Z(t_{\tilde{\beta}})$	-2.76	-2.73	-2.73	-2.73	-2.73
$Z(\Phi_2)$	3.84	3.78	3.79	3.79	3.76
$Z(t_{\alpha^*})$	0.23	0.23	0.24	0.24	0.25
$Z(t_{\mu^*})$	-0.31	-0.31	-0.32	-0.32	-0.33
$Z(\Phi_1)$	2.01	2.01	2.00	2.00	1.99
$Z(t_{\hat{\alpha}})$	-1.91	-1.91	-1.91	-1.90	-1.88

Critical Values

Phillips-Perron Tests.	5%	1%
$Z(\Phi_3)$	6.34	8.43
$Z(t_{\tilde{\alpha}})$	-3.43	-3.99
$Z(t_{\tilde{\beta}})$	2.79	3.49
$Z(\Phi_2)$	4.75	6.22
$Z(t_{\alpha^*})$	-2.88	-3.46
$Z(t_{\mu^*})$	2.53	3.19
$Z(\Phi_1)$	4.63	6.52
$Z(t_{\hat{\alpha}})$	-1.95	-2.85

TABLE 2: TESTS FOR COINTEGRATION BETWEEN FORWARD AND SPOT RATES.

A) One-Month Forward Maturity

	U. K.	Germany	France	Swiss.	Japan
$\hat{Z}_{\alpha}$ std	-46.60[a]	-55.14[a]	-54.62[a]	-53.00[a]	-41.85[a]
$\hat{Z}_{\alpha}$ μ	-46.56[a]	-55.02[a]	-54.61[a]	-53.63[a]	-41.85[a]
$\hat{Z}_{\alpha}$ $\mu\beta$	-53.51[a]	-56.00[a]	-56.53[a]	-55.44[a]	-42.93[a]
$\hat{Z}_{t}$ std	-5.02[a]	-5.50[a]	-5.46[a]	-5.36[a]	-4.74[a]
$\hat{Z}_{t}$ μ	-5.02[a]	-5.50[a]	-5.46[a]	-5.36[a]	-4.74[a]
$\hat{Z}_{t}$ $\mu\beta$	-5.35[a]	-5.60[a]	-5.57[a]	-5.50[a]	-4.81[a]

B) Two-Month Forward Maturity

	U. K.	Germany	France	Swiss.	Japan
$\hat{Z}_{\alpha}$ std	-20.26[b]	-24.82[b]	-25.55[b]	-23.26[b]	-19.34[b]
$\hat{Z}_{\alpha}$ μ	-20.02	-24.71[b]	-25.45[b]	-23.17[b]	-19.32
$\hat{Z}_{\alpha}$ $\mu\beta$	-25.59	-26.83	-27.46[b]	-26.18	-19.92
$\hat{Z}_{t}$ std	-3.10[b]	-3.55[a]	-3.64[a]	-3.36[b]	-3.19[b]
$\hat{Z}_{t}$ μ	-3.08	-3.54[b]	-3.64[b]	-3.35	-3.18
$\hat{Z}_{t}$ $\mu\beta$	-3.57[a]	-3.72	-3.79	-3.62	-3.19

C) Three-Month Forward Maturity

	U. K.	Germany	France	Swiss.	Japan
$\hat{Z}_{\alpha}$ std	-12.89	-15.52	-16.10[b]	-13.81	-10.05
$\hat{Z}_{\alpha}$ μ	-12.76	-15.49	-16.06	-13.75	-9.97
$\hat{Z}_{\alpha}$ $\mu\beta$	-17.95	-16.98	-17.46	-16.08	-10.57
$\hat{Z}_{t}$ std	-2.47	-2.88	-2.94[b]	-2.66	-2.32
$\hat{Z}_{t}$ μ	-2.46	-2.87	-2.94	-2.66	-2.31
$\hat{Z}_{t}$ $\mu\beta$	-3.00	-2.99	-3.04	-2.88	-2.34

TABLE 2 (concluded).

D) Six-Month Forward Maturity					
	U. K.	Germany	France	Swiss.	Japan
$\hat{Z}_\alpha$ std	-4.27	-5.69	-6.44	-5.03	-5.33
$\hat{Z}_\alpha$ μ	-4.26	-5.54	-6.37	-4.29	-5.21
$\hat{Z}_\alpha$ $\mu\beta$	-11.92	-7.31	-8.33	-7.07	-3.17
$\hat{Z}_t$ std	-1.36	-1.74	-1.84	-1.64	-1.82
$\hat{Z}_t$ μ	-1.31	-1.73	-1.83	-1.62	-1.81
$\hat{Z}_t$ $\mu\beta$	-2.07	-1.89	-2.04	-1.87	-1.14

Note: [a] and [b] denote significance at the 1% and 5% level, respectively.

Phillips-Ouliaris Critical Values

	1%	5%
$\hat{Z}_\alpha$ std	-22.83	-15.64
$\hat{Z}_\alpha$ μ	-28.32	-20.49
$\hat{Z}_\alpha$ $\mu\beta$	-35.42	-27.09
$\hat{Z}_t$ std	-3.39	-2.76
$\hat{Z}_t$ μ	-3.96	-3.36
$\hat{Z}_t$ $\mu\beta$	-4.36	-3.80

Note: std denotes standard, μ denotes demeaned, and $\mu\beta$ denotes demeaned and detrended.

TABLE 3: PHILLIPS-HANSEN FULLY MODIFIED ESTIMATES FOR ONE-MONTH FORWARD MATURITY

$$s^{*}_{t+k}=\alpha+\beta f^{k}_{t}+\varepsilon_{t+k}$$

	U. K.	Germany	France	Swiss.	Japan
α_{OLS}	0.0118	-0.0191	-0.0229	-0.0160	-0.2967
α_{FM}	0.0129	-0.0186	-0.0218	-0.0150	-0.1958
	(0.0087)	(0.0115)	(0.0296)	(0.0098)	(0.0611)
β_{OLS}	0.9826	1.0144	1.0071	1.0156	1.0043
β_{FM}	0.9791	1.0140	1.0065	1.0145	1.0025
	(0.0241)	(0.0126)	(0.0146)	(0.0135)	(0.0116)
R^2_{OLS}	0.8872	0.9684	0.9586	0.9643	0.9755
R^2_{FM}	0.8872	0.9684	0.9586	0.9643	0.9755
$W_{\alpha}(1)$	2.1994	2.6114	0.5434	2.2355	0.1026
$W_{\beta}(1)$	0.7492	1.2184	0.2058	1.1491	0.0460

Notes: [a] and [b] denote significance at the 1% and 5% level, respectively.
Standard errors in parenthesis.

TABLE 4: PHILLIPS-HANSEN FULLY MODIFIED ESTIMATES FOR TWO-MONTH FORWARD MATURITY

$$s^{+}_{t+k}=\alpha+\beta f^{k}_{t}+\varepsilon_{t+k}$$

	U. K.	Germany	France	Swiss.	Japan
α_{OLS}	0.0325	-0.0342	-0.0377	-0.0272	-0.0537
α_{FM}	0.0336	-0.3475	-0.0400	-0.0274	-0.0377
	(0.0129)	(0.0160)	(0.0407)	(0.0140)	(0.0932)
β_{OLS}	0.9359	1.0247	1.0101	1.0247	1.0073
β_{FM}	0.9323	1.0252	1.0101	1.0249	1.0044
	(0.0364)	(0.0176)	(0.0198)	(0.0193)	(0.0176)
R^2_{OLS}	0.7621	0.9426	0.9262	0.9322	0.9465
R^2_{FM}	0.7621	0.9426	0.9262	0.9322	0.9465
$W_{\alpha}(1)$	6.7848[a]	4.4688[b]	0.9183	3.8259	0.1635
$W_{\beta}(1)$	3.4647	2.0568	0.2899	1.6699	0.0627

Notes: [a] and [b] denote significance at the 1% and 5% level, respectively.
Standard errors in parenthesis.