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Foreign exchange market efficiency and common stochastic trends*

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In a recent paper, Baillie and Bollerslev (1989), using daily data from 1980 to 1985, identified six common stochastic trends in a vector of seven nominal exchange rates implying the existence of one cointegrating vector. Cointegration implies that (Granger) causality must run in at least one direction, that is, at least one of the exchange rates is predictable using current available information. This result has been interpreted as foreign exchange market inefficiency, by many. Another interpretation is suggested if the stationary linear combination of spot rates proxies for a time varying risk premium in some way. Then these results could be explained in a rational and risk averse market. This possibility is eliminated if the time series properties of the risk premium are incompatible with those of the error correction term. Specifically, it is demonstrated that the forward risk premium is non-stationary for the exchange rates that comprise the exchange rate cointegration relationship. In this paper, the existence of common stochastic trends in a vector of nominal exchange rates is tested over the period 1974 to 1991. The efficiency of foreign exchange markets is then tested by examining the implications of stochastic trends in the forward premium and what this means for the time series properties of a time-varying forward risk premium. (JEL F31, G14).

In a recent paper, Baillie and Bollerslev (1989), using daily data from 1980 to 1985, identified six common stochastic trends in a vector of seven nominal exchange rates implying the existence of one cointegrating vector. This had very interesting implications for foreign market efficiency since cointegration implies that (Granger) causality must run in at least one direction. That is, at least one of the future exchange rate changes is predictable using current available information. This result has usually been interpreted as evidence that foreign exchange markets are inefficient.[1] But this is not the only interpretation possible. The stationary linear combination of spot rates that comprises the cointegrating relation may proxy for a stationary and time varying forward risk premium in some way. The empirical refutation of the forward rate unbiasedness hypothesis itself implies significant predictability of future spot rate changes. This result does

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not necessarily imply foreign exchange market inefficiency if agents are risk averse. The theorized existence of a time varying forward risk premium has come to be widely cited as a reasonable explanation of the forward rate bias. If the predictability implied by cointegration among different spot exchange rates is simply capturing the predictability implied by the forward risk premium, we can reconcile Baillie and Bollerslev's results with conditions for efficient markets under risk aversion.

Other researchers have examined the evidence of common stochastic trends among different exchange rates. These include, but are not limited to, Hakkio and Rush (1989), MacDonald and Taylor (1989), Coleman (1990), Sephton and Larson (1991), and Copeland (1991). Hakkio and Rush (1989) examine monthly data from 1975 to 1986 on the British pound and Deutsche mark exchange rates *vis à vis* the US dollar and find no evidence of cointegration. They interpret this as evidence of foreign exchange market efficiency. Sephton and Larsen (1991) demonstrate that the evidence of common stochastic trends in a system of four exchange rates, is dependent upon the time period being tested. They find that using data from 1975 to 1986, they cannot reject one cointegrating vector. The evidence of common stochastic trends among different exchange rates is fragile, at best.

In order to motivate the methodology, the existence of common stochastic trends in a vector of nominal exchange rates is tested over the period 1974 to 1991. Like most previous studies, the evidence presented here finds weak evidence of cointegration among different spot rates. The efficiency of foreign exchange markets is then tested by examining the existence of stochastic trends (*i.e.* unit roots) in the forward premium and what this means for the time series properties of a time-varying risk premium. The evidence supports the hypothesis of a unit root in the forward premium making it *impossible* for the stationary error correction term (which is a linear combination of lagged spot rates), to be serving as an instrument for the forward risk premium.[2] This leads to the conclusion that foreign exchange markets violate the condition for weak form efficiency, barring any other plausible explanation for the predictability of exchange rate changes implied by cointegration among different exchange rates.

The rest of the paper is organized as follows. Section I provides the results for a system of three bilateral exchange rates relative to the US dollar. Section II discusses the implications of the results from Section I, with respect to foreign exchange market efficiency. Evidence of a non-stationary risk premium in the forward rate is provided. This result, coupled with the results of Section I, provide a basis for rejecting efficiency in the foreign exchange market. Section III provides a summary of results and conclusions.

I. Common trends in nominal exchange rates

The existence of common stochastic trends among exchange rates has been examined by Hakkio and Rush (1989), MacDonald and Taylor (1989), Coleman (1990), Sephton and Larson (1991), and Copeland (1991), inter alia, but the results are mixed. In an effort to motivate the subsequent tests for market efficiency, cointegration analysis is applied to a vector of nominal exchange rates. The spot and 30-day forward exchange rates used in this study are the British pound, German Deutsche mark, and Canadian dollar, all relative to the US dollar. The data are monthly dollars per unit of foreign currency, sampled at the close of trading on the last business day of each month from January 1974 to December 1991. The data are from the Data Resources, Inc. data tape and have been converted to natural logarithms.

The estimation of common stochastic trends is conducted using the method introduced by Johansen (1988). This methodology is, by now, well known. So I will direct the uninitiated reader to Johansen (1988) and Baillie and Bollerslev (1989) for a more thorough discussion of the methodology and its application to exchange rates across countries.

When there exists cointegration among a vector of variables Engle and Granger (1987) have demonstrated that the proper specification for estimation and hypothesis testing is in a vector error correction model (VECM). The VECM describes the adjustment of the variables in the system to deviations from the steady state equilibrium implied by cointegration. The existence of an error correction parameter entering significantly into an equation of the VECM implies causality (predictability) from the error correction term (which is a linear combination of the *past* levels of the variables) to the dependent variable in that equation. It is this result that has been loosely interpreted as evidence of inefficiency in the foreign exchange market. Cointegration itself implies that at least one such relationship must exist among the variables, in this case spot exchange rates across different countries.

A starting point for the determination of common stochastic trends is the examination of the univariate properties of the data over the sample period. Specifically, a necessary condition for the existence of cointegration is that all the variables be integrated of the same order. The results of univariate unit root tests are presented in Table 1.

The test statistics in Table 1 are those associated with the augmented Dickey–Fuller (ADF) regression (see Said and Dickey (1984)). The τ_μ statistic is the simple t-ratio and the ρ_μ statistic is the normalized bias. Schwert (1987, 1989) has done extensive Monte Carlo studies of the small sample properties of

TABLE 1. Unit root tests.

Series	Levels τ_μ	Levels ρ_μ	Diffs τ_μ	Diffs ρ_μ
S_t^{UK}	−2.11	−4.34	−5.72*	−131.87*
S_t^{WG}	−1.07	−2.21	−5.74*	−138.04*
S_t^{CA}	−2.13	−3.08	−5.72*	−162.51*
F_t^{UK}	−2.12	−4.39	−5.74*	−131.74*
F_t^{WG}	−1.09	−2.25	−5.73*	−137.14*
F_t^{CA}	−2.15	−3.11	−5.76*	−163.94*

*Significant at the 5% level. 5% critical values are −2.89 and −14.2 for the τ_μ and ρ_μ tests respectively.
[a]ADF tests with $k = 4$.

unit root tests under different error structures. His results indicate that the ADF tests have better power, and so these are employed using critical values supplied in Schwert (1987).

The results of Table 1 are based upon an ADF lag truncated at $k=4$. The results are robust to a wide choice of lag length, *i.e.* from 0 to 15, and lead to the fairly non-controversial conclusion that nominal exchange rates (and forward rates) possess a unit root in the AR polynomial representation in levels. Their first differences are stationary.[3]

The existence of cointegration among different exchange rates was tested using the Johansen methodology, where the vector of I(1) variables is $X_t = [S_t^{UK}, S_t^{WG}, S_t^{CA}]'$. The results of the cointegration analysis are presented in Table 2. The column labeled H_0: represents the null hypotheses being tested in each row. The test for the number of cointegration vectors is carried out sequentially. The first null tested is that of zero cointegration vectors. This null is rejected at the 5 percent level.[4] The null hypotheses of no more than one and no more than two cointegration vectors cannot be rejected.

Table 2 shows evidence of one cointegrating vector among three variables, or two common trends driving three series. This result is consistent with the studies by Baillie and Bollerslev (1989) and Sephton and Larson (1991). The lag length of the chosen VAR was five. This lag length was chosen on the basis of the Box–Ljung Q-statistic for serial correlation in the residuals (shown in column 4 of Table 3). The lag was chosen such that the errors were reduced to white noise statistically, as suggested by Johansen (1988). Table 3 provides estimates of the cointegration vector, β, and the error correction coefficients, α, where β has been

TABLE 2. Exchange rate cointegration results[a].

H_0:	Trace	Max eigen	5% Trace[b]	5% Max eigen[b]
$r=0$	36.41	26.80	29.68	20.97
$r \leq 1$	9.61	7.57	15.41	14.07
$r \leq 2$	2.04	2.04	3.76	3.76

[a]Lag length for VAR is set at 5.
[b]Critical values are taken from Table 1 in Osterwald–Lenum (1990).

TABLE 3. Cointegration statistics and diagnostics.

Equation	β'	s.e.[a]	$\alpha \times 100$	B-L Q-stat[b]
S_t^{UK}	1.000*	0.49	4.952	3.52
S_t^{WG}	2.959*	0.83	−2.613	2.33
S_t^{CA}	−4.267*	1.61	3.725*	14.84

*Significant at 5% level.
[a]Asymptotic standard errors.
[b]Box–Ljung test, distributed Chi-Square(12).

TABLE 4. VECM summary statistics.

Equation	$\alpha \times 100$	t-stat	F-UK[a]	F-GER[a]	F-CAN[a]
S_t^{UK}	4.95	1.87	3.90*	0.69	1.38
S_t^{WG}	−2.61	−0.90	0.29	2.73*	0.82
S_t^{CA}	3.73*	3.41	0.56	0.80	0.94

* Significant at the 5% level.
[a] F-test of joint significance of lagged changes of variable j in equation i.

normalized such that the element corresponding to the pound/dollar exchange rate is equal to one.[5] The asymptotic standard errors are computed using the square root of the $\chi^2(1)$ test statistic of the null that $\hat{\beta}_j = 0$, given in Johansen (1988), as an asymptotic Student t-test.

The interpretation of foreign exchange market inefficiency is a result of the significant error correction coefficient, α, on the Canadian dollar spot rate. It can be shown that this result implies that Canadian dollar spot rate depreciation is predictable, *i.e.* Granger-caused, by past values of the Canadian dollar, British pound, and German Deutsche mark spot rates. This result need not, however, imply the Canadian dollar/US dollar foreign exchange market is inefficient if there exists a *stationary* time varying risk premium in that market. The risk premium must be stationary since the cointegrating vector is stationary by definition. Therefore, the only way the cointegrating vector could possibly be serving as an instrument or proxy for the risk premium is when the risk premium has compatible time series properties.

Further evidence is presented in Table 4 that suggests this predictability or causality is only relevant to the Canadian dollar exchange rate. The concepts of causality and exogeneity are important in the context of the system of exchange rates examined here. The issue of exogeneity/causality is examined in Table 4 in more detail. These statistics were obtained by estimating the VECM by SUR, using the super-consistently estimated cointegration relations as the error correction variables. An autocorrelation and heteroskedasticity consistent covariance matrix is estimated using the methodology proposed by Newey and West (1987) giving asymptotically correct test statistics. The F-tests shown are tests of the joint significance of the lagged ΔX_t's in each equation. These are the traditional Granger–Sims causality tests and can be used to determine if the dependent variable is strongly exogenous, given initial acceptance of weak exogeneity (*i.e.* an insignificant error correction coefficient).

The evidence supports a one-way causal link, from the stationary linear combination of spot rates to the Canadian dollar/US dollar exchange rate (although the pound error correction coefficient is marginally significant). None of the other-variable lagged changes are jointly significant in any equation. These results imply that the pound/dollar and Deutsche mark/dollar exchange rates are strongly exogenous to the system, and that they Granger-cause Canadian dollar/US dollar exchange rate changes. The two exogenous exchange rates can be given the economic interpretation as the underlying common stochastic trends in the trivariate system (see Park, 1990 and Gonzalo and Granger, 1991). Temporary disequilibriums, as evidenced by non-zero values of the error correction term, elicit no response from the pound and Deutsche mark exchange rates. Only the Canadian dollar rate adjusts to bring the system back into equilibrium. Therefore, changes in the Canadian dollar rate are predictable based on past disequilibriums. This is the result that most have interpreted as market inefficiency.[6]

II. Market efficiency issues

The existence of a time varying but covariance stationary risk premium in the foreign exchange market could in fact explain the predictability found in the foreign exchange market, implied by the results of Section II. In this section, the forward premium (discount) on foreign exchange is decomposed into spot rate depreciation, the rational expectations error term, and the forward risk premium. This decomposition highlights the time series properties of spot and forward exchange rates necessary to produce a stationary risk premium. If these necessary properties have no support empirically, then neither does the 'proxy' hypothesis introduced above.

It is generally believed that the violation of forward rate unbiasedness is due to the existence of risk aversion on the part of foreign exchange market participants. Thus, they require a premium on forward contracts that expose them to exchange rate risk. If the error correction term in the VECM $\beta' X_t$, is highly correlated with the risk premium, it may be serving as an instrument (or proxy) for the risk premium in the foreign exchange market. But since the error correction term is stationary by definition, the risk premium must also be stationary if this explanation is to be a valid one. If, on the other hand, the risk premium is a non-stationary series, it cannot be correlated with the stationary error correction term of the spot rate system analyzed above. Therefore, the necessary condition of risk premium, and thus forward premium (discount), stationarity is examined.

Consider equation ⟨1⟩, that relates the forward rate to the future spot rate (all variables in natural log form), a risk premium, δ, and a rational expectations forecast error, ε.

$$\langle 1\rangle \qquad F_t = S_{t+1} + \delta_{t+1} + \varepsilon_{t+1}.$$

The future spot rate can be further decomposed into today's spot rate plus future depreciation as in equation ⟨2⟩.

$$\langle 2\rangle \qquad S_{t+1} = S_t + \Delta S_{t+1}.$$

By substituting equation ⟨2⟩ into ⟨1⟩, we get,

$$\langle 3\rangle \qquad F_t - S_t = \Delta S_{t+1} + \delta_{t+1} + \varepsilon_{t+1},$$

which shows the forward premium as in the sum of spot rate depreciation, risk premium and rational expectations error. If the risk premium is covariance stationary, then the forward and spot exchange rates must be cointegrated (have one common trend) with the cointegrating vector of [1, −1].[7]

In a recent paper, Crowder (1992b) has demonstrated evidence of a unit root in the forward premium of four exchange rates using daily data over the period 1982 to 1992. This result is shown to be not only supported empirically but theoretically as well. Consider the evidence provided by Meese and Rogoff (1988) on the existence of unit roots in the differentials between US and foreign interest rates. By imposing the stronger and more meaningful efficiency condition (see Dwyer and Wallace, 1992) of covered interest arbitrage, the existence of a unit root in the interest differential implies a unit root in the forward premium. If there does exist a stochastic trend in the interest differential, as Meese and Rogoff suggest, then cointegration of spot and forward rates, *i.e.* stationary forward premiums, would imply positive risk free arbitrage profits with probability of one!

Figure 1 displays plots of the forward premiums for the three exchange rates being examined. Although each series displays periods of persistence, it is sometimes difficult to judge the order of integration simply based on a plot of the data, therefore the sample autocorrelations for each forward premium series are computed. Evidence of the non-stationarity of the forward premiums is provided in Figure 2, where the sample autocorrelation function is plotted for each forward premium, along with 5 percent confidence intervals. Each plot displays the tell-tale signs of non-stationarity, a smooth and slow rate of decay. It is important to point out that it is not the absolute size of the autocorrelations that suggests non-stationarity, only their significance at long lags and slow rate of decay.

To provide a more formal test of whether the risk premium of the exchange rates under examination here are covariance stationary (or alternatively if the spot and forward rate are cointegrated with $\beta = [1, -1]$), unit root tests are performed on the forward premium as defined in equation ⟨3⟩, *i.e.* $F_t - S_t$. Table 5 presents results of those tests.

The tests are the ADF τ_μ test analogous to those presented for the levels of the spot and forward rates in Table 1.[8] Lag length on each series was determined by using the Box–Ljung Q-statistic to test for residual serial correlation of up to twelfth order. This test is distributed $\chi^2(12)$ yielding a 5 percent critical value of 21.03. The evidence supports the null that the forward premium is non-stationary

TABLE 5. Unit root tests on forward premium[a].

Forward premium	Lag length	ADF τ_μ test	B-L Q-stat[b]
FP_t^{UK}	4	−3.14*	18.89
FP_t^{WG}	11	−0.06	18.79
FP_t^{CA}	5	−2.86	19.09

* Denotes significance at the 5% level.
[a]Forward premium is defined as $F_t - S_t$.
[b]Test of serial correlation in the regression residuals and is distributed $\chi^2(12)$.

for two of the three exchange rates. The lag length necessary to yield uncorrelated errors for each test is 4, 11 and 5 for the pound, Deutsche mark and Canadian dollar forward premiums, respectively. Based on this criterion for choosing lag length, the pound forward premium appears stationary, but the Deutsche mark and Canadian dollar forward premiums appear non-stationary. Remember that

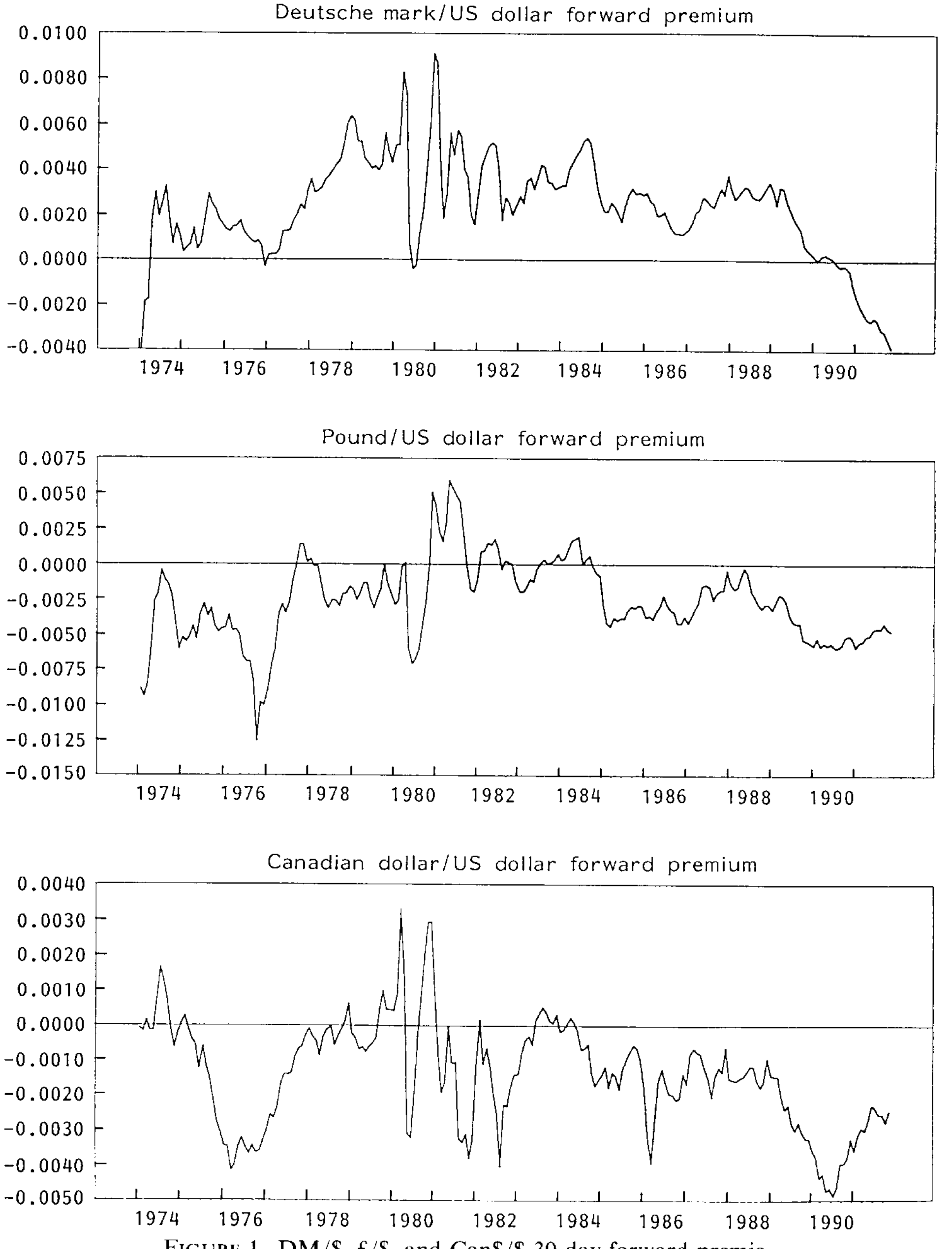

FIGURE 1. DM/$, £/$, and Can$/$ 30-day forward premia.

it is the Canadian dollar forward premium that is of particular interest, since it is this exchange rate that appears to be predictable, at least in part, by past disequilibriums among the pound, Deutsche mark and Canadian dollar exchange rates.

Finally, it should be noted that although the Box–Ljung statistic revealed no

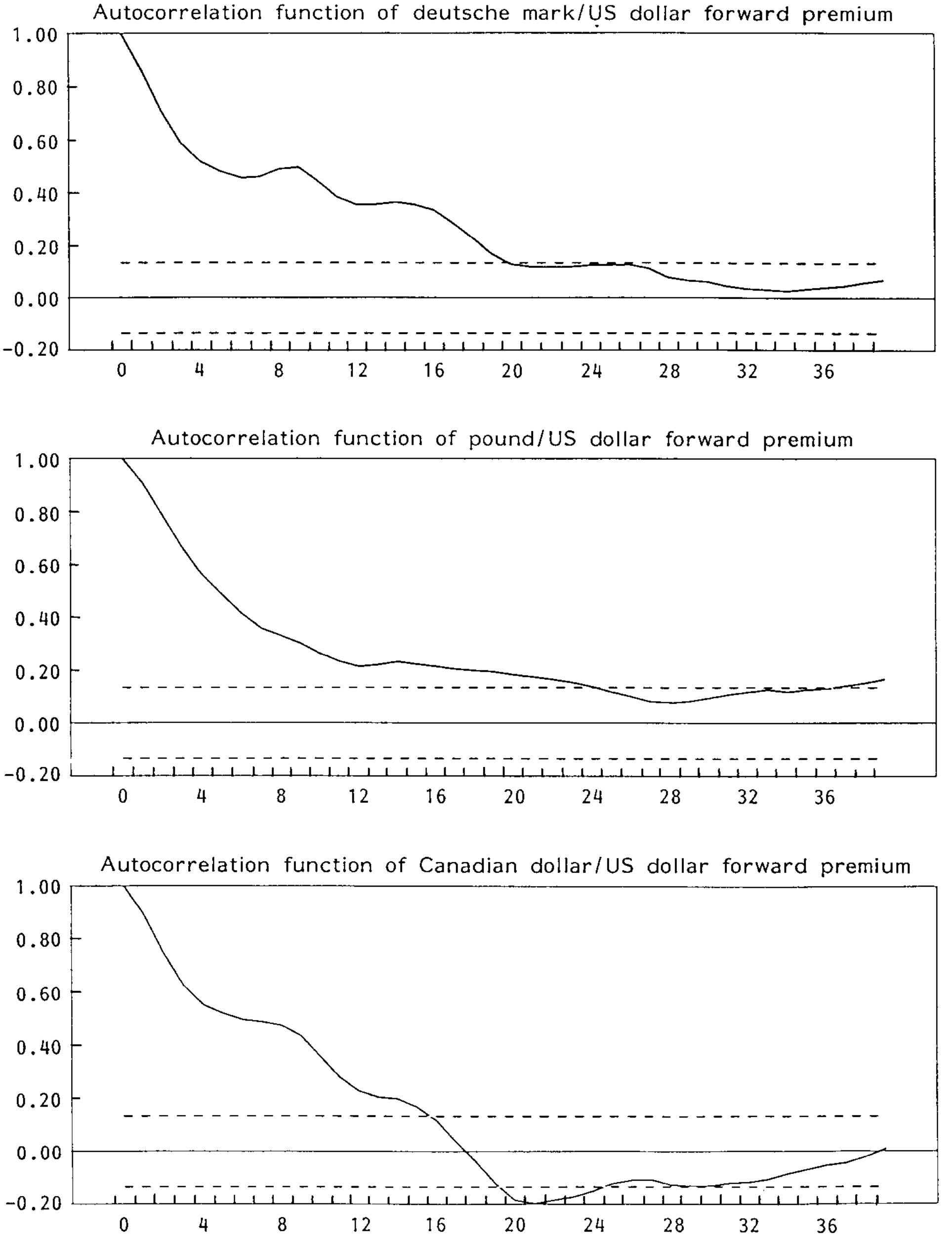

FIGURE 2. Correlagrams of the DM/$, £/$, and Can$/$ 30-day forward premia.

significant serial correlation in the residuals from the pound ADF equation after four lags, the seventh lag in the ADF equation is significant at high levels. At this lag length (7), the null of a unit root in the pound/dollar forward premium cannot be rejected.[9] Therefore, the evidence on this series is still quite inconclusive.

There has recently been criticism of standard unit root tests that take non-stationarity as the null and require strong evidence to the contrary in order to reject. To address this issue, Kwiatkowski *et al.* (KPSS) (1992) have developed a test for stationarity based on the LM score from a regression of the variable on a constant and possibly a time trend. The test is calculated as,

$$\langle 4 \rangle \qquad \hat{\eta}_\mu = \frac{\sum_{t=1}^{T} S_t^2}{T^2 s^2(k)},$$

where S_t^2 is the partial sum process of the residuals from the regression described and $s^2(k)$ is a consistent estimate of the error variance using the procedure advocated by Newey and West (1987) that is based on a Bartlett window adjustment using the first k sample autocovariances. KPSS derive critical values of 0.463 and 0.347 for the 5 percent and 10 percent levels of significance, respectively. The KPSS test was performed for each forward premium series using $k = 0$, 4, 8 and 12. The results of these tests are presented in Table 6.

The results of the KPSS tests for level stationarity reject the null for all three forward premiums at lag truncations of length zero and four at the 5 percent level of significance. The stationary null is also rejected for the three forward premiums at least at the 10 percent level when the lag truncation is selected to be eight. The null cannot be rejected for the pound/dollar forward premium when the Bartlett window lag length is set at twelve, but the null is rejected for the Deutsche mark/dollar and Canadian dollar/dollar forward premiums at the 10 percent level of significance. KPSS provide power and size properties of these tests under varying lag truncations using Monte Carlo simulations. They suggest, as a compromise between $k = 4$ (where size distortions are significant) and $k = 12$ (where power is substantially decreased), that k be set equal to eight in order to

TABLE 6. KPSS tests on forward premium[a].

Bartlett window lag truncation	FP_t^{UK}	FP_t^{WG}	FP_t^{CA}
0	2.22*	3.66*	3.65*
4	0.54*	0.92*	0.90*
8	0.35**	0.59*	0.58*
12	0.28	0.45**	0.45**

* Denotes significance at the 5% level. ** Denotes significance at the 10% level.
[a]Forward premium is defined as $F_t - S_t$.

TABLE 7. Forward and spot rate cointegration results[a].

Johansen VECM lag length	UK $H_0{:}r=0$	UK $H_0{:}r\leqslant 1$	WG $H_0{:}r=0$	WG $H_0{:}r\leqslant 1$	CA $H_0{:}r=0$	CA $H_0;r\leqslant 1$
1	29.14*	6.63	22.60*	0.30	27.19*	2.20
2	24.87*	5.13	19.47*	1.26	22.72*	3.13
3	20.64*	4.68	15.01	0.78	19.67*	2.96
4	18.28*	4.70	10.35	0.69	15.92	3.58
5	14.93	3.82	6.73	0.30	14.45	3.26

* Denotes significance at the 5% level. See Table 2 notes for critical values.
[a] $X_t=[F_t, S_t]'$.

obtain reasonable inference. Based on this suggestion, all forward premiums reject the null of level stationarity, at least at the 10 percent level of significance.

Several studies have presented evidence to the effect that spot and forward rates are cointegrated implying that the risk premium must be stationary.[10] To demonstrate the weakness of these previous results, Johansen cointegration tests were conducted on the spot and 30-day forward rates for each exchange rate, *i.e.* $X_t=[F_t, S_t]'$. The cointegration results in Table 7 are for the VECM with lags 1 through 6.

It is clear from Table 7 that the evidence of cointegration between the spot and 30-day forward rates of the same exchange rate is critically dependent on the lag length chosen. Unlike the results for cointegration among different exchanges rates presented in Table 2, the evidence in favor of cointegration weakens as lags are added to the VECM. This result could be due to large size distortions in the tests at low lag lengths or large power deterioration at higher lags. But, this result is also consistent with a unit root in the forward risk premium. The evidence casts considerable doubt on the hypothesis that spot and forward exchange rates are cointegrated with $\beta=[1, -1]'$.

The results of this section lead to the conclusion that evidence of cointegration between spot and one month forward rates of the same exchange rate is weak at best. Further evidence on the time series properties of the forward premiums leads to the conclusion that at least the Deutsche mark/dollar and Canadian dollar/dollar forward premiums have a unit root in the autoregressive polynomial representation.[11] Given the assumptions laid out earlier, this implies that the risk premium is non-stationary. Therefore the error correction term, which is stationary by definition, could not be serving as a proxy for the risk premium, due to their differing orders of integration. This evidence suggests that the predictability implied by common trends in the spot foreign exchanges rates analyzed in Section II is consistent with a violation of the conditions for market efficiency. Other likely explanations are based on the poor small sample properties of unit root estimates. It is widely understood that unit root tests have very low power against stationary alternatives. Therefore, it may be that forward premiums are stationary but have very long memory. This would be consistent with market efficiency and cointegration among different spot rates. Baillie and Bollerslev (1992) present evidence that the error correction term among different spot rates is fractionally integrated. Their evidence is able to reconcile cointegration among different exchange rates if the forward premium is also well represented by a fractionally integrated series.

III. Conclusions

Previous studies of common stochastic trends in a system of nominal exchange rates rejected non-cointegration. This finding implies a long run equilibrium relationship in the system of spot exchange rates, that itself implies causality of at least one of the exchange rates in the system. This causality has been interpreted by most as inefficiency in the foreign exchange market. Another interpretation is possible, though, if the error correction term from the VECM, is serving as an instrument for a stationary and time-varying risk premium.

This study demonstrates the existence of two common stochastic trends driving three nominal exchange rates, the British pound, German Deutsche mark, and Canadian dollar all vis-à-vis the US dollar. The question of whether this constitutes market inefficiency or not was tested by examining the time series properties of the forward premium. The presence of a unit root in the forward premium could not be rejected, implying that the forward risk premium is integrated of order one, given the assumption of rational expectations. This evidence leads to the conclusion that the presence of common stochastic trends in the system of exchange rates could not be serving as an instrument for the forward risk premium. Foreign exchange market efficiency may hold if the forward premium is a long memory stationary process, such as a fractionally integrated series. Research efforts in this direction may help in reconciling market efficiency and common trends among different exchange rates.

Notes

1. This is the interpretation given by Hakkio and Rush (1989) of Baillie and Bollerslev's (1989) result.
2. It is important that the terms forward premium, meaning the premium or discount at which future dollars are selling relative to spot dollars, and forward risk premium, which is the premium required by risk averse agents above expected spot rate depreciation, not be confused.
3. For a review of issues surrounding unit roots in macroeconomic data, see Campbell and Perron (1991). Note that pretesting the variables for a unit root is not a requirement for the cointegration techniques used here, since the tests can discriminate between stationary and non-stationary series.
4. It should be noted that the tests for cointegration are quite sensitive to the choice of lag in the VAR. To examine robustness, the tests were computed for VAR lag lengths of 1 to 8. The null of zero cointegration vectors is not rejected at lags 1 to 4, but is rejected at lags 5 to 8. Johansen (1988) suggests examining the stability of the cointegration vector estimates in order to determine if a long-run equilibrium relationship does in fact exist. In this regard all estimates of the first column of β are very similar. Still, if the linear combination of spot rates does not have substantial serial correlation, the use of longer lags is not warranted.

5. Germany's participation in the ERM and the EMS's target zones may account for the Deutsche mark and pound being cointegrated in the late 1980s and early 1990s. But the Canadian dollar enters the cointegration vector significantly, implying that all of the cointegration evidence cannot be explained by the ERM or EMS.
6. Recently, Baillie and Bollerslev (1992) have presented evidence that the cointegrating vector among different exchange rates may be fractionally integrated. This would be consistent with the mixed evidence on the existence of cointegration among exchange rates. It is the case, however, that fractionally integrated time series are virtually indistinguishable from ARIMA(0,1,1) models with large negative MA errors. Crowder (1992b) provides clear evidence that the data generating process of the forward premium is entirely consistent with an ARIMA(0,1,1) with MA parameters on the order of -0.5 to -0.9. It is unlikely that this issue is to be resolved in favor of one representation over another, *i.e.* ARIMA vs. ARFIMA. In terms of this study, the important point is that using compatible methodologies, it is demonstrated that the time series properties of the forward premium are inconsistent with those of exchange rate cointegration. One is non-stationary and the other is stationary.
7. The spot rate depreciation is stationary based on results from Table 1. The rational expectations forecast error is assumed stationary, since a unit root in the forecast error implies irrationality. Therefore, by assumption, the only term left over that could be causing non-stationarity in the forward premium is the risk premium term.
8. The normalized bias tests yielded virtually the same inference and are therefore excluded to economize on space.
9. Unit root tests of up to twelve lags were computed. Results are available upon request.
10. Baillie and Bollerslev (1989) present such findings.
11. Box–Jenkins ARIMA analysis was conducted on each of the three forward premiums. The models chosen by this analysis were ARIMA(0,1,1) for the UK and Canadian forward premiums and ARIMA(0,1,2) for the German forward premium. In all three models, the MA(1) parameter estimate (the sum of the MA(1) and MA(2) parameter estimates for Germany) was close to -0.9. Even so, Wald tests of parameter redundancy, *i.e.* MA unit root, all rejected at high levels. Although the properties of this test statistic are unknown in the presence of MA unit roots, the results are consistent with those given in Table 6 which essentially tests the same thing.

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