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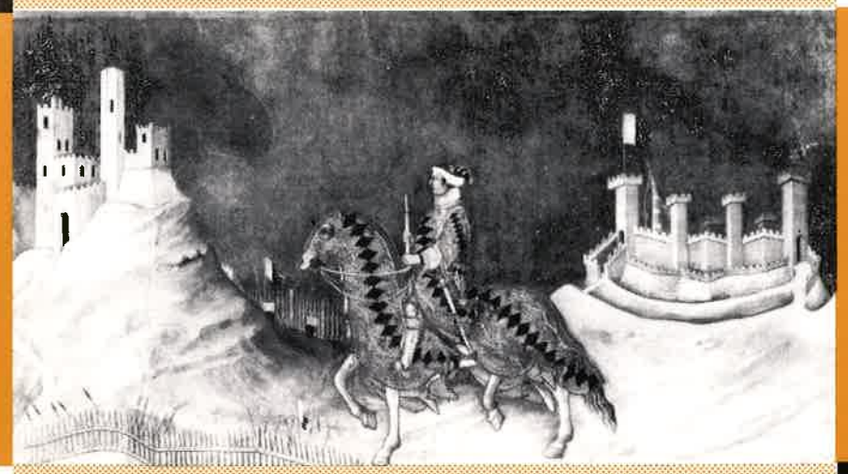
QUADERNI DEL DIPARTIMENTO
DI ECONOMIA POLITICA

Giulio Cifarelli

COINTEGRATION AND THE UNBIASED EFFICIENCY OF THE FORWARD EXCHANGE RATE



QUADERNI DEL DIPARTIMENTO DI ECONOMIA POLITICA

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Giulio Cifarelli

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Siena, novembre 1990

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I. Introduction*

Exchange rate market efficiency has been tested with the help of relationships which, for the same set of data, provide conflicting results. An accurate analysis of the modelling procedure is therefore of paramount importance. In this paper we provide a possible interpretation of these findings with the help of cointegration analysis and show that the apparent contradictions are due to the differences between the short run and the long run properties of the exchange rate time series.

II. Empirical Tests of Exchange Rate Market Efficiency

In an efficient exchange market investors are assumed to be risk neutral and to derive rational expectations. The covered and uncovered interest parity conditions suggest that in such a market the forward rate be regarded as the

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expected future spot rate. More precisely, the forward exchange rate is said to be an unbiased and efficient expectation of the future spot exchange rate if

1) $E(s_t|\psi_{t-1}) - f_{t-1} = 0$,

where E(.) is the expectations operator, ψ_{t-1} is the information set of economic agents at time t-1, s_t is the logarithm of the spot exchange rate at time t and f_{t-1} is the logarithm of the corresponding one period ahead forward exchange rate, set forth in period t-1. Tests of the unbiased efficiency (of the forward rate) are assumed to be tests of exchange rate market efficiency.

Equation 1) can be tested empirically by estimating equation 2):

2) $s_t = \alpha + \beta f_{t-1} + u_{2t}$.

The unbiased efficiency hypothesis is assumed to hold if $\hat{\alpha} = 0$, $\hat{\beta} = 1$ and u_{2t} is a white noise error term.[1] The forward exchange rate is seen then as the optimal predictor of the future spot rate, in the sense that it minimises the mean squared error. Empirical estimates of equation 2), performed by B. Cornell (1977), J.A. Frenkel (1977,1979,1980), R. Macdonald (1983), R.T. Baillie, R.E. Lippens and P.C. McMahon (1983) and others, tend to support the unbiased efficiency hypothesis. Equation 2) has been the object of various criticisms, due to the alleged nonstationarity of the exchange rate time series.[2]

J.F.O. Bilson (1981), R.E. Cumby and M. Obstfeld (1984), D. Blake, M. Beenstock and V. Brasse (1986), K.A. Froot and J.A.

Frenkel (1989) and others estimate a variant of equation 2), in which exchange rate changes are explained by the forward premium of the previous period:

3) $(s_t - s_{t-1}) = \alpha' + \beta'(f_{t-1} - s_{t-1}) + u_{3t}$.

The time series involved are now stationary, and, here too, the unbiased efficiency hypothesis is assumed to hold if $\hat{\alpha}' = 0$, $\hat{\beta}' = 1$ and u_{3t} is a white noise error process. Empirical estimates of equation 3) tend to contradict the unbiased efficiency hypothesis.[3] The same set of data, estimated with equations 2) and 3) will often provide contradictory results: market efficiency holds if tested with equation 2) and is rejected if tested with equation 3). In this paper we try to reconcile these contradictory results with the help of cointegration analysis.

III. Cointegration Analysis of the Exchange Rate Time Series

We investigate the behaviour of the exchange rate between the italian lira, the german mark, the U.S. dollar and the british pound from march 1973 to april 1985, with the help of monthly non overlapping data.[4] As usual in cointegration analysis, we pretest the relevant time series for unit roots. We apply - following D. Corbae and S. Ouliaris (1986), R.T. Baillie and T. Bollerlsev (1989) and P.C. McMahon (1989) - the univariate unit root test statistics set forth by P.C.B. Phillips (1987) and P.C.B. Phillips and P. Perron (1988).[5] The statistics set forth in table 1 are those required to implement the test strategy suggested by P.

TABLE 1

1973.3 1985.4

	$Z(\Phi_3)^*$	$Z(\Phi_2)^*$	$Z(\Phi_1)^*$	$Z(t_{\hat{\alpha}})^*$	$Z(t_{\hat{\alpha}_c})^*$
s_t					
Lit\L	3.33	3.01	1.53	-2.58	-0.76
DM\L	3.35	3.13	4.43	-2.18	-2.47
S\L	0.88	1.73	1.72	-1.06	-0.14
f_{t-1}					
Lit\L	2.97	2.71	1.63	-2.43	-0.97
DM\L	4.23	3.97	5.55	-2.50	-2.76
S\L	1.15	1.99	1.95	-0.73	0.46
Δs_t					
Lit\L	62.32	41.55	62.23	-11.16	-11.16
DM\L	60.92	40.64	60.18	-11.04	-10.97
S\L	57.88	38.59	56.50	-10.75	-10.63
Δf_{t-1}					
Lit\L	73.41	48.94	73.39	-12.11	-12.12
DM\L	62.71	41.81	60.84	-11.20	-11.03
S\L	64.11	42.74	62.56	-11.32	-11.19
$(f_{t-1} - s_{t-1})$					
Lit\L	62.32	41.55	62.23	-11.16	-11.16
DM\L	60.92	40.64	60.18	-11.04	-10.97
S\L	57.88	38.59	56.50	-10.75	-10.63

*THE NUMBER OF AUTOCORRELATIONS IN THE RESIDUAL SUM VARIANCE ESTIMATOR IS 3. THE EXACT FORMULAS OF THE Z STATISTICS ARE SET FORTH IN THE APPENDIX.

CRITICAL VALUES AT THE 5% LEVEL OF SIGNIFICANCE : $Z(\Phi_3) = 6.49$; $Z(\Phi_2) = 4.88$; $Z(\Phi_1) = 4.71$; $Z(t_{\hat{\alpha}}) = -3.45$; $Z(t_{\hat{\alpha}_c}) = -2.89$.

Perron (1988).[6] They strongly suggest that s_t and f_{t-1} are I(1) and that Δs_t , Δf_{t-1} and $(f_{t-1} - s_{t-1})$ are I(0).[7] These findings also suggest that the s_t and f_{t-1} time series are cointegrated, i.e. that there is a long run (equilibrium) relationship between them, quantified by the cointegration vector. An obvious choice for the cointegrating relationship is provided by equation 2). The OLS estimates of this equation and the corresponding relevant statistics of the standard tests of cointegration are set forth in table 2. If s_t and f_{t-1} are I(1) and if the residuals of equation 2), $\hat{u}_{2t}$, are I(0), then s_t and f_{t-1} will be cointegrated of order 1.[8] The empirical findings

TABLE 2

1973.3 1985.4

	α^*	β^*	$\bar{R}^2$	CRDW	$Z(\bar{\alpha})^{**}$	$Z(t_{\hat{\alpha}_c})^{**}$	ADF
Lit\L	0.13	0.983	0.98	1.83	-135.82	-11.12	-5.61
DM\L	0.05	0.966	0.97	1.78	-126.59	-10.81	-5.51
S\L	-0.05	1.002	0.98	1.80	-130.76	-10.50	-5.81

* WE DO NOT REPORT THE COEFFICIENT t RATIOS, WHICH ARE BIASED.

**THE NUMBER OF AUTOCORRELATIONS IN THE RESIDUAL SUM VARIANCE ESTIMATOR IS 3. THE EXACT FORMULAS OF THE Z STATISTICS ARE SET FORTH IN THE APPENDIX.

CRITICAL VALUES AT THE 5% LEVEL OF SIGNIFICANCE : $Z(\bar{\alpha}) = -20.49$; ADF, $Z(t_{\hat{\alpha}_c}) = -3.36$; CRDW = 0.39.

strongly suggest that $\hat{u}_{2t}$ is I(0) and so reject the null hypothesis of no cointegration between the s_t and f_{t-1} time series.

IV. The Error Correction Representation

The Granger representation theorem (C.W.J. Granger (1983)) asserts that if a set of variables are cointegrated, then there always exists an error correcting formulation of the dynamic model which describes their behaviour over time and vice versa. Following the two-step procedure suggested by R.F. Engle and C.W.J. Granger (1987), the residuals from the cointegrating equation - equation 2) - have been used as an error correcting variable in the following unrestricted autoregression:

$$4)\ \Delta s_t = a + \sum_{i=1}^{8} b_i \Delta s_{t-i} + \sum_{i=0}^{8} c_i \Delta f_{t-1-i} + d\hat{u}_{2t-1} + e_t.$$

Following the "general to simple" modelling procedure, we obtained the parsimonious representations of the data generating processes of the exchange rates of interest set forth in table 3. The empirical findings are highly satisfactory: the parsimonious versions of equation 4) are all data acceptable (on an F test) and exhibit no sign of either autoregressive or moving average structure in the residuals of up to order 6 (LM tests). The estimates are stable over time and have - with one exception - good out of sample forecasting performances. In order to assess the relevance of the two-step estimation procedure, we have relaxed - following S.G. Hall (1986) and T.J. Jenkinson (1986) - the restrictions implied by the cointegrating regression, and have included the one period lagged values of s_t and f_{t-1} unrestrictedly in the dynamic models of table 3. The quality of fit is not affected, a result which supports the

TABLE 3

ERROR CORRECTING MODELS 1973.4 1985.4

Lit\L : 4.1) $\Delta s_t = a + b_4 \Delta s_{t-4} + c_0 \Delta f_{t-1} + d\hat{u}_{2t-1} + e_t$

a	Δs_{t-4}	Δf_{t-1}	$\hat{u}_{2t-1}$	$\bar{R}^2$	S.E.	D.W.	LM(6)	C(4,137)	AR(1)	F(15,126)	H(8,141)
0.01	-0.21	-1.27	1.23	0.07	0.0274	1.98	3.58	0.85	0.19	0.93	1.98
(3.17)	(-1.92)	(-2.14)	(2.16)								

	s_{t-1}	f_{t-2}						C(5,135)		H(8,140)
-0.02	-0.21	-1.59	1.55	-1.54	0.08	0.0272	1.99	3.61	1.14	0.86	1.91
(-0.26)	(-1.90)	(-2.48)	(2.50)	(-2.52)							

DM\L : 4.2) $\Delta s_t = b_1 \Delta s_{t-1} + d\hat{u}_{2t-1} + e_t$

Δs_{t-1}	$\hat{u}_{2t-1}$	$\bar{R}^2$	S.E.	D.W.	LM(6)	C(2,141)	AR(1)	F(17,126)	H(8,143)
0.82	-0.74	0.02	0.0301	1.99	3.83	0.54	2.38	1.17	0.83
(2.31)	(-2.17)								

	s_{t-1}	f_{t-2}	a				C(4,137)		H(8,141)
-1.17	1.21	-1.25	0.04	0.04	0.0298	2.00	4.15	1.04	3.89	0.91
(-1.32)	(1.44)	(-1.48)	(2.20)							

S\L : 4.3) $\Delta s_t = a + b_1 \Delta s_{t-1} + c_5 \Delta f_{t-5} + d\hat{u}_{2t-1} + e_t$

a	Δs_{t-1}	Δf_{t-5}	$\hat{u}_{2t-1}$	$\bar{R}^2$	S.E.	D.W.	LM(6)	C(4,137)	AR(1)	F(15,126)	H(8,141)
-0.01	-1.58	0.19	1.63	0.05	0.0291	1.98	2.98	0.62	2.73	0.89	3.10*
(-2.64)	(-2.19)	(2.20)	(2.37)								

	s_{t-1}	f_{t-2}						C(5,135)		H(8,140)
0.002	-1.62	0.20	1.67	-1.69	0.04	0.0291	1.99	3.34	0.66	2.35	2.34*
(0.23)	(-2.25)	(2.33)	(2.44)	(-2.44)							

* : NULL HYPOTHESIS REJECTED AT THE 5% LEVEL.

LM(.) : LAGRANGE MULTIPLIER TEST FOR 6TH-ORDER ERROR STRUCTURE.[9]

C(.) : CHOW TEST FOR PARAMETER STABILITY IN THE 1973.4 1979.2, 1979.3 1985.4 TIME PERIODS. AR(.) : TEST FOR FIRST ORDER ARCH.

F(.) : F TEST OF THE PARSIMONIOUS REPRESENTATION RESTRICTIONS.

H(.) : CHOW TEST OF POST-SAMPLE PREDICTIVE ACCURACY.

t RATIOS ARE COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

cointegration hypothesis. These findings have some controversial economic implications. The cointegrating regression (equation 2)) can be seen as a long run equilibrium relationship. s_t and f_{t-1} follow a common trend, they may drift far apart in the short run, but in the long run market forces will bring them together. If $\hat{\alpha} = 0$, $\hat{\beta} = 1$ and $\hat{U}_{2t}$ is white noise, then the forward exchange rate could well be seen - always in the long run - as an unbiased and efficient predictor of the future spot rate. These properties do not have to hold also in the short run. Indeed, the estimates of the error correcting models set forth in table 3 show that the lagged equilibrium error ($\hat{U}_{2t-1}$) plays a relevant role in determining short run exchange rate behaviour. This finding can be interpreted - as sugested by R.T. Baillie and T. Bollerslev (1989) - as a violation of market efficiency and/or as evidence of a time varying risk premium.[10] The deviations from unbiasedness have to be stationary or transitory, however, since s_t and f_{t-1} are cointegrated. We have thus to introduce a short run versus long run distinction in the theory of exchange rate market efficiency.

V. The Error Correction Specification and the Tests of Unbiased Efficiency

As pointed out above, estimating models in the levels when the (predetermined) variables are nonstationary invalidates the standard asymptotic theory. As a consequence, the conventional tests of the efficiency hypothesis based on the estimates of equation 2) are incorrect. Cointegration between the s_t and f_{t-1} time series invalidates also the more

sophisticated tests suggested by C.S. Hakkio (1981), R.T. Baillie, R.E. Lippens and P.C. McMahon (1983) and others, tests based on the estimation of vector autoregressions of first differences of spot and forward exchange rates. Indeed - as pointed out by R.F. Engle and C.W.J. Granger (1987) - models of this kind are misspecified, since the error correction term is missing (too many unit roots are imposed on the data).

Surprisingly, the simple tests of the unbiased efficiency hypothesis based on the estimation of equation 3) may well be correct, even if the s_t and f_{t-1} time series are nonstationary and cointegrated. Indeed, equations 2) and 3) may be complementary: if equation 2) is the cointegrating regression, equation 3) is a restrictive version of the corresponding error correcting model , equation 4). If $\hat{\beta} = 1$ in the cointegrating equation 2), equation 3) can be rewritten as the following error correcting equation:

5) $\Delta s_t = a + c_0\Delta f_{t-1} - c_0(s_{t-1} - \hat{\alpha} - f_{t-2}) + u_{5t}$.

Equation 5) is a data acceptable parsimonious version of equation 4) above if $b_1 = b_2 = \ldots = b_8 = 0$, $c_1 = c_2 = \ldots = c_8 = 0$ and $c_0 + d = 0$. In order to assess the empirical relevance of equation 3) we have rewritten equation 4) as follows,

4') $\Delta s_t = a + \sum_{i=1}^{8} b_i\Delta s_{t-i} + \sum_{i=0}^{8} c_i\Delta f_{t-1-i} + d(s_{t-1} - \hat{\alpha} - f_{t-2}) -$

$d(\hat{\beta} - 1)f_{t-2} + e_t$

and have tested the validity of the restrictions $b_1 = b_2 = \ldots = b_8 = c_1 = c_2 = \ldots = c_8 = 0$, $c_0 + d = 0$ and $d(\hat{\beta} - 1) = 0$. Only if

these restrictions are rejected by the data is equation 3) misspecified. These tests and the estimates of equation 3) are set forth in table 4.

TABLE 4

1973.4 1985.4

3) $\Delta S_t = \alpha' + \beta'(F_{t-1} - S_{t-1}) + u_{3t}$.

	α'	β'	$F_{H'}$	R^2	S.E.	D.W.	LM(6)	C(2,141)##	AR(1)	H(9,143)
Lit\i	0.01	-1.41	1.046	0.06	0.0276	2.07	7.71	1.12	1.37	1.56
	(3.16)	(-2.40)								
DM\i	-0.01	-1.50	1.543	0.02	0.0303	1.80	4.33	4.57*	8.49*	0.82
	(-2.35)	(-1.60)								
S\i	-0.01	-1.71	1.044	0.04	0.0294	1.87	8.55	1.83	4.66*	2.14*
	(-2.88)	(-2.46)								

* : NULL HYPOTHESIS REJECTED AT RHE 5% LEVEL.

AR(.) : TEST FOR FIRST ORDER ARCH.

$F_{H'}$: F TEST OF THE H' HYPOTHESIS THAT IN EQUATION 4) $b_1 = \ldots = b_8 = c_1 = \ldots = c_8 = 0$; $c_0 + d = 0$ AND $d(\hat{\beta}-1) = 0$. AT THE 5% LEVEL OF SIGNIFICANCE, F(18,125) = 1.64.

C(.) : CHOW TEST FOR PARAMETER STABILITY IN THE 1973.4 1979.2, 1979.3 1985.4 TIME PERIODS.

H(.) : CHOW TEST OF POST-SAMPLE PREDICTIVE ACCURACY.

THE t RATIOS ARE COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

The quality of fit is rather good and we cannot reject the restrictions of interest at the 5% level of significance. The error correcting models set forth in table 3, however, are more stable and provide a closer approximation of the true

data generating processes, as shown by the outcome of the Davidson MacKinnon J tests between non nested models set forth in table 5 below. Indeed, the fitted values of the error

TABLE 5

1973.4 1985.4

	$t_{\Delta\hat{S}_t}$*	$t_{\Delta\hat{S}_t}$**
Lit\i	1.75	2.37
DM\i	1.71	2.62
S\i	1.18	2.25

*t RATIO OF THE COEFFICIENT OF THE FITTED VALUE OF EQUATION 3) IN THE ESTIMATION OF THE ERROR CORRECTING EQUATION OF TABLE 3.

**t RATIO OF THE COEFFICIENT OF THE FITTED VALUE OF THE ERROR CORRECTING EQUATION OF TABLE 3 IN THE ESTIMATION OF EQUATION 3).

correcting equations are significant if added as explanatory variables to the estimates of equation 3), whereas the opposite is not the case.

VI. Conclusion

We have shown in this paper that the apparent contradiction between the results obtained with two popular tests of the unbiased efficiency hypothesis (i.e. of market efficiency) are due to the difference between the short run and the long run properties of the data, and can be explained with the help of cointegration analysis. One of the empirical relationships is a (long run) cointegrating regression and the

other is a restricted version of the corresponding error correcting model.

The economic implications of these findings, however, are not clearcut, since the theory of efficient markets does not accommodate well to the short run versus long run distinction.

Appendix

The test statistics used in this paper - set forth by P.C.B. Phillips (1987), P.C.B. Phillips and P. Perron (1988) and P.C.B. Phillips and S. Ouliaris (1990) - are defined as follows.

For equation 1*), $y_t = \hat{\mu}_t + \hat{\alpha} y_{t-1} + \hat{u}_t$, we have:

A.1) $$Z(t_{\hat{\alpha}}) = (\hat{s}/s_{Tl}) t_{\hat{\alpha}} - (1/2 s_{Tl})(s_{Tl}^2 - \hat{s}^2)\left[T^{-2} \sum_{t=1}^{T} (y_{t-1} - \bar{Y}_{-1})^2 \right]^{-1/2}$$

A.2) $$Z(\Phi_1) = (\hat{s}^2/s_{Tl}^2)\Phi_1 - (1/2 s_{Tl}^2)(s_{Tl}^2 - \hat{s}^2)\Big\{ T(\hat{\alpha} - 1) - 1/4(s_{Tl}^2 - \hat{s}^2)$$

$$\left[T^{-2} \sum_{t=1}^{T} (y_{t-1} - \bar{Y}_{-1})^2 \right]^{-1} \Big\}$$

with

$$\Phi_1 = (2\hat{s}^2)^{-1} T(s^2 - \hat{s}^2).$$

The variance $\hat{s}^2$, the standard t ratio for $\hat{\alpha} = 1$ $t_{\hat{\alpha}}$, the residual variance under the appopriate null hypothesis s^2 and the mean of $\bar{Y}$ are defined in the following way:

$$\hat{s}^2 = T^{-1} \sum_{t=1}^{T} (y - \hat{\mu} - \hat{\alpha} y_{t-1})^2$$

$$t_{\hat{\alpha}} = (\hat{\alpha} - \alpha) \left\{ \sum_{t=1}^{T} (y_{t-1} - \bar{Y}_{-1})^2 \right\}^{1/2} / \hat{s}$$

$$s^2 = T^{-1} \sum_{t=1}^{T} (y_t - y_{t-1})^2.$$

s_{Tl}^2 is a consistent estimator of the variance of the sum of the residuals under the null hypothesis and is analysed below.

For equation 2*), $y_t = \tilde{\mu} + \tilde{\beta}(t - T/2) + \tilde{\alpha} y_{t-1} + \tilde{u}_t$, we have:

A.3) $$Z(t_{\tilde{\alpha}}) = (\tilde{s}/s_{Tl}) t_{\tilde{\alpha}} - (T^3/4\sqrt{3} D_x^{1/2} s_{Tl})(s_{Tl}^2 - \tilde{s}^2)$$

A.4) $$Z(\Phi_2) = (\tilde{s}^2/s_{Tl}^2)\Phi_2 - (1/3 s_{Tl}^2)(s_{Tl}^2 - \tilde{s}^2)[\, T(\tilde{\alpha} - 1) - (T^6/48 D_x)$$

$$(s_{Tl}^2 - \tilde{s}^2) \,]$$

A.5) $$Z(\Phi_3) = (\tilde{s}^2/s_{Tl}^2)\Phi_3 - (1/2 s_{Tl}^2)(s_{Tl}^2 - \tilde{s}^2)[\, T(\tilde{\alpha} - 1) - (T^6/48 D_x)$$

$$(s_{Tl}^2 - \tilde{s}^2) \,],$$

with

$$\Phi_2 = (3\tilde{s}^2)^{-1} T(s^2 - \tilde{s}^2)$$

$$\Phi_3 = (2\tilde{s}^2)^{-1} T[\, s^2 - (\bar{Y}_0 - \bar{Y}_{-1})^2 - \tilde{s}^2 \,]$$

$$t_{\tilde{\alpha}} = (\tilde{\alpha} - \alpha) / (\tilde{s}^2 C_3)^{1/2}$$

$$\tilde{s}^2 = T^{-1} \sum_{t=1}^{T} [\, y_t - \tilde{\mu} - \tilde{\beta}(t - T/2) - \tilde{\alpha} y_{t-1} \,]^2$$

where C_i is the (i, i) element of the determinant D_x of the squared matrix of the regressors of equation 2*).

For equation 3*), $y_t = \bar{\alpha} y_{t-1} + \bar{u}_t$, we have:

A.6) $$Z(\bar{\alpha}) = T(\bar{\alpha} - 1) - (1/2)(s_{Tl}^2 - \bar{s}^2) \left[T^{-2} \sum_{t=1}^{T} y_{t-1}^2 \right]^{-1}$$

A.7) $Z(t_{\tilde{\alpha}}) = (\tilde{s}/s_{T,l})t_{\tilde{\alpha}} - (1/2s_{T,l})(s^2_{T,l} - \tilde{s}^2)[\, T^{-2} \sum_{t=1}^{T} y^2_{t-1}\,]^{-1/2}$,

where

$$t_{\tilde{\alpha}} = (\tilde{\alpha} - \alpha)(\sum_{t=1}^{T} y^2_{t-1})^{1/2} / \tilde{s}$$

$$\tilde{s} = T^{-1} \sum_{t=1}^{T} (y_t - \tilde{\alpha} y_{t-1})^2 .$$

In the expressions above the residual sum variance estimator is defined as

$$s^2_{T,l} = T^{-1} \sum_{t=1}^{T} x^2_t + 2T^{-1} \sum_{\tau=1}^{l} \omega_{\tau,l} \sum_{t=\tau+1}^{T} x_t x_{t-\tau} ,$$

where $x_t = \hat{u}_t$, $\tilde{u}_t$ or $\bar{u}_t$, and $\omega_{\tau,l} = 1+\tau/(l+1)$. l - the truncation lag - corresponds to the maximum number of autocorrelations in the disturbances. We have computed the Z statistics with l = 0, 1, 2, 3, 4. The results obtained with l = 3 are reported in tables 1 and 2.

Notes

(1) The forward rate is assumed to be an unbiased forecast of the future spot rate if $\hat{\alpha} = 0$ and $\hat{\beta} = 1$. This forecast is also efficient if the residual contains no price sensitive information, such that $E(u_{2t}u_{2t+i}) = 0 \;\forall\; i > 0$.

(2) As pointed out by R.A. Meese and K.J. Singleton (1982), if the f_{t-1} time series is nonstationary, the conventional tests of the efficiency hypothesis in equation 2) above are incorrect since the asymptotic distribution theory used for inference can no longer be applied.

(3) The forward premium is thus a biased predictor of the

future spot exchange rate change. E. Fama (1984), R. Hodrick and S. Srivastava (1984) and D. Hsieh (1984) attribute this result to a relevant and time varying risk premium. D. Longworth (1981), J.F.O. Bilson (1981) and K.A. Froot and J.A. Frenkel (1989) interpret this finding as evidence of systematic expectational errors, i.e. as a failure of the rational expectations hypothesis.

(4) We use end of month data published on the "Financial Times". The time period under examination starts with the currency crisis of January-February 1973 (dollar devaluation, E.M.S. crisis) and ends in April 1985, on the eve of a realignment of the currencies belonging to the E.M.S.. The lira dropped out of the E.M.S. in January-February 1973 and reentered it in March 1979. The german mark has never left the E.M.S.. In the time period under examination both the dollar and the pound continued to float freely.

(5) The standard Dickey-Fuller procedure (D.A. Dickey and W.A. Fuller (1979,1891)) can be applied if the error terms of the test relationships are independently and identically distributed. The conditions imposed on the error terms by the Perron-Phillips tests used in this paper are much less restrictive and allow for many weakly dependent and heterogeneously distributed innovations. For a clear exposition of this point, see C. Giannini and R. Mosconi (1989).

(6) The statistics set forth above are based on the OLS estimation of the following relationships:

$1^*)$ $y_t = \hat{\mu} + \hat{\alpha} y_{t-1} + \hat{u}_t$

$2^*)$ $y_t = \tilde{\mu} + \tilde{\beta}(t - T/2) + \tilde{\alpha} y_{t-1} + \tilde{u}_t$,

where T is the number of observations in the sample. The residuals $\hat{u}_t$ and $\tilde{u}_t$ are quite general and may be stationary

ARMA processes, with a time varying variance. The $Z(\Phi_2)$ and $Z(\Phi_3)$ F statistics deal with the following hypotheses:

$Z(\Phi_2)$ $\quad H_1^0 : (\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (0, 0, 1)$ in equation 2^*)

$Z(\Phi_3)$ $\quad H_2^0 : (\tilde{\mu}, \tilde{\beta}, \tilde{\alpha}) = (\tilde{\mu}, 0, 1)$ in equation 2^*).

The $Z(t_{\tilde{\alpha}})$ t statistic is used to test the following hypothesis:

$H_3^0 : \tilde{\alpha} = 1$ in equation 2^*).

P. Perron suggests that we start with the estimation of equation 2^*) and assess whether the hypotheses H_3^0 and H_2^0 can be rejected with the help of the statistics $Z(t_{\tilde{\alpha}})$ and $Z(\Phi_3)$. If these hypotheses cannot be rejected the analysis of the estimates of equation 1^*) becomes relevant. The (more powerful) statistics derived from these estimates are not invariant with respect to the drift coefficient μ. We have therefore to compute $Z(\Phi_2)$. If - as is the case with the s_t and f_{t-1} time series analysed in table 1 - the $Z(\Phi_2)$ statistic suggests that the hypothesis H_1^0 cannot be rejected, the use of the statistics derived from equation 1^*) becomes appropriate. The following unit root hypotheses are then assessed using the $Z(\Phi_1)$ and $Z(t_{\hat{\alpha}})$ statistics:

$Z(\Phi_1)$ $\quad H_4^0 : (\hat{\mu}, \hat{\alpha}) = (0, 1)$ in equation 1^*)

$Z(t_{\hat{\alpha}})$ $\quad H_5^0 : \hat{\alpha} = 1$ in equation 1^*).

Under their respective null hypotheses, $Z(\Phi_3)$, $Z(\Phi_2)$, $Z(\Phi_1)$, $Z(t_{\tilde{\alpha}})$ and $Z(t_{\hat{\alpha}})$ will have the Φ_3, Φ_2, Φ_1, τ_τ and τ_μ Dickey-Fuller distributions. (W.A. Fuller (1976), D.A. Dickey and W.A. Fuller (1981)).

(7) For the Δs_t, Δf_{t-1} and $(f_{t-1} - s_{t-1})$ time series the null hypothesis can be rejected also against the explosive alternative ($\hat{\alpha}$, $\tilde{\alpha} > 1$ in equations 1^*) and 2^*)), since the estimated values of the $Z(t_{\hat{\alpha}})$ and $Z(t_{\tilde{\alpha}})$ statistics lie on the left hand tails of their respective Dickey-Fuller distribution.

(8) We test for the presence of unit roots in the $\hat{u}_{2t}$ time series. The CRDW statistic is the D.W. statistic of the cointegrating regression. Under the null of no cointegration $\hat{u}_{2t}$ contains an unit root and CRDW is not significantly different from zero. The ADF statistic is the t statistic of the coefficient γ in the following regression:

$$\Delta\hat{u}_{2t} = \gamma\hat{u}_{2t-1} + \sum_{i=1}^{4} \lambda_i \Delta\hat{u}_{2t-i} + \epsilon_t.$$

The critical values of these statistics have been tabulated by R.F. Engle and B.S. Yoo (1987) with 100 observations. The $Z(\tilde{\alpha})$ and $Z(t_{\tilde{\alpha}})$ statistics are used to assess the relevance of the hypothesis that $\tilde{\alpha} = 1$ in the OLS estimates of equation 3^*),

3^*) $y_t = \tilde{\alpha}y_{t-1} + \tilde{u}_t$ $\quad$ (where $y_t = \hat{u}_{2t}$).

There is no intercept in equation 3^*) since - by construction - the OLS residuals $\hat{u}_{2t}$ have mean zero. These statistics have the same properties, in this context, of the Z statistics examined above. Their critical values have been tabulated by P.C.B. Phillips and S. Ouliaris (1990) with 500 observations.

(9) The Lagrange multiplier test for sixth order (MA or AR) error structure LM(6) is distributed asymptotically as χ_6^2 on the null. The Chow test for parameter constancy C(k,T-2k) is distributed as F(k,T-2k) on the null, where T is the number of observations in the sample and k is the number of regressors. AR(1) is asymptotically distributed as χ_1^2 on the null of no ARCH (Autoregressive Conditional Heteroskedasticity) effect. The Chow test H(n,T-k)) is a measure goodness of fit over n post-sample time periods and is distributed as F(n,T-k) on the null of constant parameters.

(10) Alternatively, the dynamics of the error correcting

models could be interpreted as evidence of the relevance of the "news" in the short run. In the "news" literature forward forecast errors ($\hat{u}_{2t}$ in this context) do not necessarily contradict exchange rate market efficiency, and are due to unanticipated changes in the factors which influence the expectations about the future behaviour of the spot exchange rate.

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