

Monetary and Asset Market Models for Sterling Exchange Rates: A Cointegration Approach

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Abstract

The aim of this paper is to investigate the determination of bilateral sterling exchange rates over the floating period 1973Q1-1990Q3. The exchange rates covered are: US dollar/pound, D-mark/pound, yen/pound, and F-franc/pound. We provide an econometric evaluation of the main exchange rate theories, using the cointegration-error correction methodology and non-nested tests. We have been unable to find any statistical evidence in support of a long-run relationship consistent with the various monetary models. The empirical results for the long-run and dynamic exchange rate equations provide strong support for a modified uncovered interest rate parity and the portfolio balance models.

I. Introduction

The purpose of this study is to model bilateral sterling exchange rates

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over the floating period 1973Q1 – 1990Q3. The exchange rates covered are: US dollar/ pound, D-mark/pound, yen/pound, and French franc/pound. We provide an econometric evaluation of the main theories of exchange rate determination, using the cointegration-error correction modeling methodology and non-nested tests.

Modelling exchange rates has attracted considerable attention over the last fifteen years, due to the crucial role exchange rates play in the interaction of quantities and prices, in macroeconomic policy, and in international finance. Past studies have concentrated primarily on US dollar exchange rates rather than on sterling rates. Those studies that have dealt with the latter have focused mostly on sterling's effective exchange rate (Beenstock *et al.* [1981], Hacche and Townend [1981], Currie and Hall [1989], Fisher *et al.* [1990]). Only Sarantis [1987] has attempted to examine the determination of the main (five) bilateral sterling exchange rates. Most significantly, past studies have tended to apply one theoretical model rather than provide an econometric comparison of rival models. There have been very few attempts to provide such an evaluation. Hacche and Townend [1981] compared the pure monetarist and Dornbusch models only. More recently, Fisher *et al.* [1990] have compared the exchange rate equations used in the main macroeconometric models of the U.K. However, the equations used in these models are basically various versions of the uncovered interest parity (UIP), so the study by Fisher *et al.* does not provide an econometric evaluation of the alternative theories of exchange rate determination.

In applying the various exchange rate models to each sterling rate, we make use of the two-stage modelling methodology proposed by Engle and Granger [1987].¹ This procedure enables us to concentrate at the first stage on testing whether there is a long-run equilibrium relationship consistent with the corresponding exchange rate model. This requires the set of variables included in each exchange rate model to form a cointegrating vector. The uniqueness of these cointegrating vectors is examined with the Johansen [1988] maximum likelihood cointegration tests. In the second

1. Although the cointegration technique has been used by some authors to examine the long-run purchasing power parity (PPP) hypothesis (Taylor [1988], Kim [1990], Sarantis and Stewart [1993]), its application to alternative exchange rate models has been limited.

stage we concentrate on modelling the short-run dynamics of those exchange rate models which are supported by the cointegration tests, using the error correction model (ECM). For a valid ECM to exist, the variables must cointegrate. Otherwise the empirical results will be subject to the Granger and Newbold [1974] problem of 'spurious' regression and might be structurally unstable. Consequently, cointegration tests might rule out some exchange rate models as inadmissible.

The paper is organized as follows. Section II outlines the alternative theories of exchange rate determination and derives corresponding equations for estimation. In Section III we investigate the integrability of time series, using both the conventional Augmented Dickey-Fuller (ADF) test and the recently proposed tests by Hylleberg *et al.* [1990] for stochastic seasonality. Section IV is concerned with uncovering cointegrating or long-run equilibrium relationships consistent with the alternative theoretical exchange rate models. In Section V we specify and estimate the short-run dynamic forms of the exchange rate equations, using an error correction presentation, and compare the rival models. The final section draws up the main conclusions.

II. Theoretical Models of Exchange Rate Determination

The theoretical literature on exchange rate determination usually focuses on a subset of equations, such as those describing the asset markets or the balance of payments, rather than a fully specified macroeconomic model. The more common models used in applied research are as follows:²

Flexible Price Monetary model (FPM): This approach combines the quantity theory of money – where fully flexible prices are determined by monetary equilibrium between a stable real money demand and nominal money supply – with strict purchasing power parity (Bilson [1978]). Assuming similar demand for money functions between two countries, the FPM model is given by

$$e_t = a_0 + a_1(m^* - m)_t - a_2(y^* - y)_t + a_3(r^* - r)_t \quad (1)$$

2. For a detailed discussion of these theories and the derivation of the relevant equations, see Baillie and McMahon [1989].

where e is the log (natural) of the exchange rate, defined as the foreign currency price of domestic currency; m , y and r represent the natural logarithms of domestic money supply, real income and nominal short term interest rate, respectively; asterisks denote the corresponding foreign variables.

Sticky Price Monetary model (SPM): The assumption of continuous price flexibility is thought to be rather unrealistic. Dornbusch [1976] has argued that in the short run prices are more likely to be sticky, due to costs of adjustment and/or lack of information. On the other hand, the money market is still assumed to adjust instantaneously, while PPP is maintained as a long-run equilibrium relationship. Frankel [1979] extended the Dornbusch model by assuming that the adjustment of the exchange rate to its equilibrium level depends on the real interest rate differential. The model assumes that PPP only holds in the long run. The SPM model can be represented by the equation (where $\pi^* - \pi$ is the expected price differential)³

$$e_t = a_0 + a_1(m^* - m)_t - a_2(y^* - y)_t - a_3(r^* - r)_t + a_4(\pi^* - \pi)_t \quad (2)$$

Hooper-Morton model (HM): Hooper and Morton [1982] extended the SPM model to allow for shifts in the real exchange rate. Such shifts were assumed to be related to movements in the current account through an expectations scheme. The derived equation (HM) can be written as (Meese and Rogoff [1983])

$$e_t = a_0 + a_1(m^* - m)_t - a_2(y^* - y)_t - a_3(r^* - r)_t + a_4(\pi^* - \pi)_t + a_5(cca)_t - a_6(cca^*)_t \quad (3)$$

where (cca) and $(cca)^*$ are the cumulated domestic and foreign current account balances respectively.

Portfolio Balance Model (PB): Most of the portfolio balance models proposed in the literature (Branson *et al.* [1977], Bisignano and Hoover [1982]) assume three assets: domestic money and bonds, and foreign assets. Sarantis [1987] extended this model to incorporate the equity market and applied it to five exchange rates against the pound with satisfactory

3. Frankel [1979] shows that this model includes as special cases a number of monetarist exchange rate models. For example, $a_3 > 0$ and $a_4 = 0$ gives the FPM model, whereas $a_3 < 0$ and $a_4 = 0$ gives the Dornbusch model.

results. Solution of Sarantis's model yields the following equation for the exchange rate:

$$e_t = e \left[\underset{+}{(p^* - p)}_t, \underset{+}{(vnso)}_t, \underset{+}{m^*}_t, \underset{-}{m}_t, \underset{\pm}{k^*}_t, \underset{\pm}{k}_t, \underset{\pm}{b^*}_t, \underset{\pm}{b}_t, \underset{-}{f^*}_t, \underset{+}{f}_t \right] \quad (4)$$

where b , k and f are the nominal stocks (in natural logarithms) of domestic bonds, equities and foreign assets (held by the private sector) respectively, asterisks denote the corresponding foreign variables, and $(vnso)$ is the value of North Sea oil.

Uncovered Interest Parity (UIP): The UIP relationship, adjusted for the influence of risk, is given by

$$e_t = \hat{e}_{t+1} - (r^* - r)_t + \gamma \chi_t \quad (5)$$

where χ is interpreted as a measure of the risk premium. The risk premium term is usually proxied by the ratio of current account to nominal GDP, (CA/NY) (Fisher *et al.* [1990]). The UIP relationship differs from the portfolio balance model (4) in assuming that domestic and foreign assets are perfect substitutes, whereas model (4) maintains the assumption of imperfect substitutability.

The UIP condition (5), whether in nominal or real form, is clearly forward-looking.⁴ The Bank of England, on the other hand, in its modelling of sterling's effective exchange rate uses a modified version of the UIP model that assumes backward-looking expectations. The derived long-run relationship is written in terms of real exchange rate (ρ_t) and real interest rate differential ($i^* - i$) and has the form⁵

$$\rho_t = \gamma_0 - \gamma_1(i^* - i)_t + \gamma_2 \chi_t \quad (6)$$

4. This model was derived by Hall [1987]. An econometric evaluation of this model (both in nominal and real form) for sterling's effective exchange rate was carried out by Fisher *et al.* [1990], using McCallum's Errors-in-Variables approach for estimating forward expectations.

5. The Bank of England uses a short-run dynamic equation which is based on the error correction model and has the form (Fisher *et al.* [1990]):

$$\Delta \rho_t = f(\Delta \rho_{t-1}, \Delta(p^* - p)_{t-4}, \Delta \chi_t, \rho_{t-1}, (i^* - i)_{t-1}, \chi_{t-1})$$

The long-run solution to this equation is given by (6). In this model exchange rate expectations are captured by past movements in the exchange rate.

where

$$\rho_t = e_t + p_t - p_t^*, \quad i_t = r_t - \pi_t, \quad i_t^* = r_t^* - \pi_t^*$$

Given the definition of the real exchange rate and real interest rate differentials, equation (6) can be written in the following semi-reduced form:

$$e_t = \alpha_0 - \alpha_1(r^* - r)_t + \alpha_2(\pi^* - \pi)_t + \alpha_3(p^* - p)_t + \alpha_4\chi_t \quad (7)$$

which implies a long-run equilibrium relationship between the nominal exchange rate, the nominal interest rate differential, the expected inflation differential, the relative price and the risk premium which we proxy by the domestic and foreign ratios of current account to GDP.

III. The Order of Integration

The stationarity of time series is usually examined by applying the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests. Recently, however, Engle *et al.* [1989], Hylleberg *et al.* [1990], and Osborn [1990] have argued that seasonally unadjusted economic time series might also have seasonal unit roots, besides the (autoregressive) unit root at zero frequency which has been the focus of attention so far in the analysis of time series (Fuller [1976], Dickey and Fuller [1981], Engle and Granger [1987]). If we ignore stochastic seasonality, the Ordinary Least Squares (OLS) estimates of the cointegrating relationships will be inconsistent (Engle *et al.* [1989]).

A seasonal economic time series $\{x_t\}$ is said to be integrated of order (d, D) , i.e. $x_t \rightarrow I(d, D)$, if the series becomes stationary after one-period differencing (unit root) and seasonal differencing D times (seasonal unit root). To test for seasonal unit roots we have applied the tests recently proposed by Hylleberg *et al.* [1990], thereafter referred to as the HEGY test. The HEGY tests are shown in Table 1. The test $\pi_1 = 0$ is for a unit root, the test $\pi_2 = 0$ is for bi-annual stochastic seasonality, and the joint test for π_3 and π_4 tests for annual stochastic seasonality. Hence, rejection of stochastic seasonality requires rejection of both a test for π_2 and a joint test for π_3 and π_4 .

The t-statistics for π_1 indicate that all variables exhibit a unit root at zero frequency (both at 1% and 5%) irrespective of the deterministic terms included in the regressions. The t-statistics for π_2 and the F-statistics for the joint

Table 1^a
HEGY Tests for Seasonal Unit Roots

USA							Germany					Japan				
Variables	Specification	$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$	LM(4)	k	$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$	LM(4)	k	$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$	LM(4)	k
e	I	-1.84	-4.96	35.38	5.94	0	-2.35	-5.14	29.27	4.50	0	-1.29	-5.53	28.1	4.61	0
	I, SD	-1.82	-4.74	34.12	5.55	0	-2.32	-4.94	30.30	7.21	0	-1.28	-5.33	27.64	4.93	0
	I, SD, T	-1.69	-4.73	33.97	7.95	0	-2.80	-5.04	30.91	3.85	0	-2.54	-5.52	28.91	2.14	0
(m^*-m)	I	-1.99	-1.72	13.07	1.00	3	-1.661	-0.692	5.906	7.338	3	-2.175	-1.086	1.467	2.738	1
	I, SD	-1.84	-2.17	13.16	1.35	3	-1.454	-2.244	13.881	6.569	3	-1.931	-1.272	12.575	3.936	0
	I, SD, T	-0.65	-2.20	12.73	2.59	3	-0.755	-2.29	13.003	6.955	3	-1.168	-2.609	7.093	4.061	0
(r^*-r)	I	-2.21	-5.06	22.53	1.67	0	-2.74	-3.59	12.66	7.20	4	-1.895	-2.757	13.749	5.149	2
	I, SD	-2.26	-5.03	22.07	2.31	0	-2.603	-4.71	31.963	6.523	1	-1.712	-2.771	17.775	4.416	2
	I, SD, T	-2.22	-4.97	21.72	2.56	0	-2.594	-4.617	30.291	6.749	1	-3.797	-4.595	44.637	4.014	0
$(\pi^*-\pi)$	I	-1.32	-5.53	31.73	1.75	0	-2.097	-5.19	27.514	4.692	0	-3.139	-5.329	24.232	1.646	0
	I, SD	-1.39	-5.65	35.12	3.78	0	-2.066	-5.007	25.281	1.177	1	-3.26	-5.138	27.615	4.433	0
	I, SD, T	-2.07	-5.68	35.78	2.97	0	-2.773	-5.234	26.67	3.933	1	-3.194	-5.037	26.006	5.783	0
$(\rho^*-\rho)$	I	-1.59	-4.49	13.13	5.75	1	-2.73	-3.47	17.74	4.60	2	-2.52	-3.21	3.50	5.04	10
	I, SD	-1.53	-4.28	13.78	4.76	1	-2.63	-3.50	20.02	6.03	2	-3.55	-3.77	12.87	3.47	8
	I, SD, T	-3.06	-2.99	11.25	2.77	3	-2.34	-3.57	20.62	9.31	2	-3.19	-4.06	13.06	6.65	8
f^*	I	3.10	-8.17	7.27	5.29	3	-0.86	-3.88	20.68	3.60	0	-0.17	-2.73	12.77	1.89	6
	I, SD	3.02	-8.06	6.17	6.93	3	-0.86	-4.15	20.55	4.82	0	-0.54	-7.47	20.42	4.71	1
	I, SD, T	3.40	-8.28	6.13	7.42	3	-1.52	-4.13	20.45	3.71	0	-4.06	-10.01	22.37	3.68	0
m^*	I	0.25	-0.55	0.68	3.79	6	-0.94	-0.15	0.76	5.98	4	-2.04	-1.24	3.08	6.20	2
	I, SD	-0.01	-4.26	42.80	2.62	0	-0.83	-2.75	15.98	6.94	2	-2.23	1.05	13.25	3.98	4
	I, SD, T	-2.26	-4.36	44.72	4.68	0	-1.68	-2.78	14.71	7.45	2	-2.91	1.10	10.75	2.51	4
$(cca)^*$	I	1.14	-1.98	6.49	5.42	9	-0.56	-1.48	2.67	1.04	8	-0.09	-1.33	4.71	5.78	4
	I, SD	1.26	-2.23	5.82	5.94	9	-0.34	-8.44	42.01	8.05	3	-0.13	-1.71	5.74	5.43	4
	I, SD, T	1.70	-2.20	5.72	5.24	9	-2.53	-8.21	62.86	7.28	1	-2.74	-1.75	5.22	0.76	4
$(CA/NY)^*$	I	-1.24	-4.74	39.06	5.30	0	-0.94	-1.11	0.69	5.78	4	-2.08	-1.04	3.71	7.31	6
	I, SD	-1.26	-4.80	37.17	5.98	0	-1.39	-4.16	36.42	2.71	0	-2.66	-2.39	5.94	6.70	1
	I, SD, T	-2.33	-4.86	38.57	1.97	0	-1.82	-4.18	36.39	2.98	0	-3.76	-2.34	5.33	4.63	1
France							United Kingdom									
Variables	Specification	$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$	LM(4)	k	Variables	Specification	$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$	LM(4)	k			
e	I	-2.11	-5.76	25.37	1.69	0	m	I	-1.81	-1.65	9.24	3.24	2			
	I, SD	-2.11	-5.56	25.64	1.64	0		I, SD	-1.68	-2.08	13.84	1.41	3			
	I, SD, T	-2.28	-5.49	24.94	0.48	0		I, SD, T	0.27	-2.48	15.54	3.29	2			
(m^*-m)	I	-1.22	-1.55	6.91	4.33	1	f	I	-0.66	-3.32	10.08	3.69	4			
	I, SD	-1.11	-4.54	33.62	3.14	0		I, SD	-0.75	-3.70	11.09	7.55	4			
	I, SD, T	-2.14	-4.62	35.29	3.26	0		I, SD, T	-0.33	-3.74	11.81	8.36	6			
(r^*-r)	I	-2.42	-5.09	36.61	3.15	0	k	I	-0.69	-2.12	10.78	4.00	4			
	I, SD	-2.49	-5.03	36.31	1.28	0		I, SD	-0.65	-2.50	16.70	3.13	4			
	I, SD, T	-2.52	-4.97	35.75	1.55	0		I, SD, T	-3.20	-1.89	10.09	8.45	5			
$(\pi^*-\pi)$	I	-1.60	-4.78	33.81	2.36	0	b	I	0.92	-2.50	4.99	4.17	4			
	I, SD	-1.62	-4.56	35.43	3.84	0		I, SD	0.90	-2.30	4.28	4.23	4			
	I, SD, T	-1.72	-4.55	35.18	3.44	0		I, SD, T	-2.51	-2.42	4.16	4.84	4			
$(\rho^*-\rho)$	I	-3.45	-4.41	7.63	6.94	9	(cca)	I	0.01	-7.65	22.40	5.00	0			
	I, SD	-3.39	-4.95	7.94	8.04	10		I, SD	-0.10	-7.06	24.14	2.42	0			
	I, SD, T	-3.49	-4.79	8.01	7.41	10		I, SD, T	1.53	-6.91	24.03	3.46	0			
f^*	I	0.60	-4.52	9.02	0.36	2	(CA/NY)	I	-1.71	-3.35	7.13	3.61	1			
	I, SD	0.62	-4.42	9.10	0.72	0		I, SD	-1.61	-3.66	11.34	4.08	1			
	I, SD, T	-1.75	-4.47	8.51	1.56	2		I, SD, T	-1.60	-3.63	11.26	4.71	1			
m^*	I	-2.04	-0.26	1.37	3.95	5	(vms)	I	-1.63	-6.32	23.92	40.6	0			
	I, SD	-2.17	-4.58	30.97	2.04	0		I, SD	-1.69	-6.34	22.62	3.02	0			
	I, SD, T	-0.20	-4.52	29.33	2.04	0		I, SD, T	-1.30	-6.21	22.00	3.07	0			
$(cca)^*$	I	1.73	-1.52	7.00	4.31	2										
	I, SD	1.69	-1.85	8.09	3.57	2										
	I, SD, T	0.57	-1.85	7.90	3.35	2										
$(CA/NY)^*$	I	-1.96	-1.11	6.87	5.68	4										
	I, SD	-2.00	-0.77	9.16	3.20	4										
	I, SD, T	-2.08	-0.75	9.08	4.53	4										

		$t: \pi_1$	$t: \pi_2$	$F: \pi_2/\pi_1$
5% Critical	I	-2.88	-1.95	3.08
Values	I, SD	-2.95	-2.94	6.57
($n = 100$)	I, SD, T	-3.53	-2.94	6.60

(a) k indicates the number of lagged dependent variables required to eliminate residual autocorrelation in the HEGY regressions; LM(4) is the Lagrange Multiplier for 4th order autocorrelation. I, SD and T are intercept, seasonal dummies, and time trend respectively used in alternative specifications of the HEGY equation.

test $\pi_3 = \pi_4$ are greater than their respective 5% critical values (which suggests rejection of the hypothesis of seasonal unit roots) for all variables, with the exception of the following:⁶

USA:	$(m^* - m), m^*, f^*, (cca)^*$
Germany:	$(m^* - m), m^*, (cca)^*, (CA/NY)^*$
Japan:	$(m^* - m), m^*, (cca)^*, (CA/NY)^*$
France:	$(m^* - m), m^*, (cca)^*, (CA/NY)^*$
UK:	m, k, b

It is interesting to note that the variables which have seasonal unit roots tend to be the same for all countries. Hylleberg *et al.* [1990] have suggested to filter out the seasonal unit roots and to test for cointegration with the filtered series. The appropriate seasonal filters are indicated by the tests. In the case of biannual seasonality the series is filtered by $(1 + L)$, whereas in the case of annual seasonality the series is filtered by $(1 + L + L^2 + L^3)$.

The t-statistics for π_1 indicate that some variables might need to be differenced more than once in order to become stationary. We have therefore used the more conventional Augmented Dickey-Fuller (ADF) test to investigate further the degree of integrability of all time series. In implementing these tests we adopted the sequential testing procedure proposed by Dickey and Pantula [1987]. This involves testing for d unit roots and working down; each time the null hypothesis of nonstationarity is rejected, d is reduced by one. The procedure stops when the null hypothesis is accepted. We initially set d equal to 2, but given the large number of variables and space constraints we only report the most relevant statistics for determining the integrability of each time series. All regressions were estimated with and without a time trend as in Dickey and Fuller [1981], and the ADF statistics for both regressions are reported in Table 2. The seasonally filtered series are used where seasonal roots were present. We notice that the degree of integration for each variable remains the same irrespective of the inclusion of the time trend; indeed the ADF statistics are remarkably close

6. Osborn [1990] also found seasonal unit roots for UK narrow money, but not for trade balance (her sample period was 1955Q1 - 1988Q4). It should also be pointed out that seasonal dummies and time trend were statistically insignificant in the overwhelming majority of variables and their inclusion made no difference to the tests.

Table 2^a
The Augmented Dickey-Fuller Test for Unit Roots

Country	Variables	I(0) ADF	ADF _t	I(1) ADF	ADF _t	I(2) ADF	ADF _t
USA	e	-1.73(1)	-1.43(1)	-6.35(0)	-6.42(0)		
	$(m^* - m)^s$	-2.20(4)	-0.71(4)	-3.16(3)	-4.06(3)		
	$(y^* - y)$	-1.17(4)	-1.71(4)	-5.85(3)	-5.82(3)		
	$(r^* - r)$	-2.73(2)	-2.71(2)	-7.63(1)	-7.60(1)		
	$(\pi^* - \pi)$	-1.35(4)	-2.03(2)	-7.01(1)	-6.96(1)		
	$(p^* - p)$	-1.59(4)	-2.36(4)	-3.72(2)	-3.86(2)		
	m^{*s}	0.26(6)	-2.66(6)	-3.08(6)	-3.56(6)		
	f^{*s}	-	-	-	-	1.71(6)	0.43(6)
	$(cca)^{*s}$	-	-	-	-	0.79(2)	-1.13(2)
	$(CA/NY)^*$	-1.04(0)	-1.90(0)	-7.89(0)	-7.83(0)		
Germany	e	-2.48(0)	-2.62(0)	-6.94(0)	-7.07(0)		
	$(m^* - m)^s$	-1.67(4)	-1.27(4)	-3.12(3)	-3.73(3)		
	$(y^* - y)$	-1.44(2)	-1.44(2)	-7.67(1)	-7.61(1)		
	$(r^* - r)$	-2.70(4)	-2.77(4)	-6.14(3)	-6.13(3)		
	$(\pi^* - \pi)$	2.42(0)	-3.12(0)	-9.49(0)	-9.68(0)		
	$(p^* - p)$	-2.73(5)	-2.30(5)	-3.17(1)	-4.01(1)		
	m^{*s}	-1.00(4)	-1.60(4)	-4.69(3)	-4.73(3)		
	f^{*s}	-1.42(1)	-2.09(1)	-7.96(1)	-7.94(1)		
	$(cca)^{*s}$	-	-	-1.64(0)	-1.58(0)	-8.65(0)	-8.56(0)
	$(CA/NY)^{*s}$	-0.89(4)	-1.44(4)	-4.62(3)	-4.63(0)		
Japan	e	-1.30(1)	-1.28(1)	-5.62(0)	-5.64(0)		
	$(m^* - m)^s$	-1.64(3)	-3.12(3)	-8.74(1)	-8.67(1)		
	$(y^* - y)$	-0.26(1)	-1.50(1)	-11.08(0)	-11.01(0)		
	$(r^* - r)$	-2.26(0)	-3.05(0)	-7.61(0)	-7.56(0)		
	$(\pi^* - \pi)$	-2.91(4)	-2.81(4)	-5.27(3)	-5.45(3)		
	$(p^* - p)$	-1.11(1)	-2.37(1)	-3.92(0)	-3.93(0)		
	m^{*s}	-1.58(5)	-2.56(5)	-3.21(3)	-3.54(3)		
	f^{*s}	-0.17(9)	-1.67(9)	-5.27(8)	-5.22(8)		
	$(cca)^{*s}$	0.19(3)	-3.11(3)	-4.06(2)	-4.16(2)		
	$(CA/NY)^{*s}$	-1.83(7)	-2.01(7)	-3.95(3)	-3.93(3)		
France	e	-2.18(1)	-2.37(1)	-7.24(0)	-7.22(0)		
	$(m^* - m)^s$	-1.24(3)	-2.80(3)	-3.69(3)	-3.65(3)		
	$(y^* - y)$	-	-	-2.26(6)	-2.31(6)	-7.58(6)	-7.50(6)
	$(r^* - r)$	-2.71(1)	-2.74(1)	-7.49(0)	-7.45(0)		
	$(\pi^* - \pi)$	-1.56(0)	-1.60(0)	-8.54(0)	-8.49(0)		
	$(p^* - p)$	-3.06(3)	-2.51(3)	-4.55(1)	-4.95(1)		
	m^{*s}	-	-	-1.93(4)	-2.78(4)	-7.16(3)	-7.11(3)
	f^{*s}	0.80(0)	-1.54(0)	-7.37(0)	-7.61(0)		
	$(cca)^{*s}$	-	-	0.28(3)	-0.51(3)	-8.01(1)	-8.38(1)
	$(CA/NY)^{*s}$	-1.96(6)	-2.04(6)	-5.92(5)	-5.87(5)		
U.K.	m^s	-1.87(4)	0.01(4)	-2.99(3)	-3.55(3)		
	f^s	-	-	3.78(3)	2.77(3)	-3.03(3)	-3.89(3)
	k^s	-0.65(6)	-3.24(6)	-3.74(5)	-3.84(5)		
	b^s	1.04(4)	-8.32(3)				
	(cca)	-	-	6.03(3)	5.92(3)	-4.68(1)	-5.58(1)
	(CA/NY)	-1.71(4)	-1.70(4)	-9.33(1)	-9.29(1)		
	$(vnso)$	-1.64(4)	-1.20(4)	-4.17(4)	-4.47(4)		

Notes: (a) ADF and ADF_t are the Augmented Dickey-Fuller tests without and with a trend. The 5% critical values for the ADF and ADF_t statistics are approximately -2.90 and -3.47 respectively (MacKinnon [1990]). The number within parenthesis is the number of lagged dependent variables required to eliminate residual autocorrelation. The superscript (s) indicates seasonally filtered variables. For the USA variables f^{*s} and $(cca)^{*s}$ it was necessary to test for I(3). For f^{*s} , the statistics are ADF = -3.25(5) and ADF_t = -3.80(5). For $(cca)^{*s}$, the statistics are ADF = -7.15(1) and ADF_t = -7.38(1).

in most cases. The broad conclusions are that all variables are clearly $I(1)$, except the following variables:

- USA: $f^s: I(3)$, $(cca)^s: I(3)$
 Germany: $(cca)^s: I(2)$
 France: $(y^* - y): I(2)$, $m^s: I(2)$, $(cca)^s: I(2)$
 UK: $f: I(2)$, $(cca): I(2)$

IV. Cointegration Relationships

The exchange rate models specified in Section II are essentially long-run equilibrium relationships derived from economic theory and their empirical validity can be investigated with the cointegration tests.⁷ Following Engle and Granger [1987] we have used the method of OLS to estimate each exchange rate model over the period 1973Q1 - 1990Q3, provided that the variables are $I(1)$, and then employed the Dickey-Fuller (DF) tests and Augmented Dickey-Fuller (ADF) tests to examine whether cointegration really exists amongst the set of variables.⁸ Whenever there is conflict between the DF and ADF tests, we use the LM autocorrelation test of the respective DF and ADF relationships (which must be free of autocorrelation) to distinguish between the two. All equations were estimated with and without seasonal dummy variables to allow for the possibility of deterministic seasonality. However, all seasonal dummies were entirely insignificant with t -values below unity and made no difference to the cointegration tests. These findings are in line with those obtained with the HEGY tests and therefore all reported regressions are those obtained without seasonal dummies.

One potential problem with the residual-based cointegration tests proposed by Engle and Granger [1987] is that in multivariate models there might be more than one cointegrating vector. Hence it is quite possible that

7. For a similar argument, see Baillie and Selover [1987] and Baillie and McMahon [1989]. An additional advantage of this modelling approach is that it provides a common framework within which we can test all these models.

8. As explained in Section III, in the case of variables with stochastic seasonality we have removed the seasonal roots using appropriate seasonal filters and it is the filtered series (which are distinguished by the subscript(s) above a variable) which are used in cointegration tests.

the equilibrium relationships uncovered by the Engle-Granger cointegration tests are a linear combination of different cointegrating vectors, so we have investigated their uniqueness by applying the Johansen [1988] multivariate cointegration tests based on a maximum likelihood procedure.

Flexible Price Monetary Model: All variables included in this exchange rate model are $I(1)$, except $(y^* - y)$ for France which is $I(2)$ and hence cannot be integrated with the exchange rate. The results reported in Table 3 strongly suggest that no cointegrating vector exists amongst this set of variables for any of the exchange rates.

Table 3^a
Cointegrating Regressions
Flexible Price Monetary Model: Equation (1)

Variables	US Dollar/ pound	D. Mark/ pound	Yen/ pound	F. Franc/ pound
Constant	-13.23 (3.41)	14.065 (4.57)	3.973 (14.64)	3.705 (9.97)
$(m^* - m)_t^s$	0.056 (0.42)	0.737 (11.33)	0.997 (5.19)	0.194 (3.66)
$(y^* - y)_t$	-3.093 (4.78)	1.030 (1.98)	-0.521 (1.18)	
$(r^* - r)_t$	2.068 (2.71)	4.391 (5.98)	2.194 (2.35)	1.662 (3.69)
\bar{R}^2	0.447	0.730	0.839	0.299
DF	-1.888	-2.290	-2.180	-2.964
[LM(4)]	[5.4]	[9.44]	[23.1]	[7.56]
ADF (k)	-2.112(2)	-2.678(4)	-2.193(4)	-2.783(1)
[LM(4)]	[5.1]	[3.6]	[5.9]	[4.06]

Notes: (a) \bar{R}^2 is the coefficient of determination adjusted for degrees of freedom; numbers in parentheses under regression coefficients are t -values; LM is the Lagrange Multiplier for 4th order residual autocorrelation (distributed as χ^2); k indicates the number of lagged dependent variables required to eliminate residual autocorrelation in the ADF regression; the 5% critical values for Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests, for $n = 50$ and $n = 100$ and for up to 5 variables, can be found in Engle and Yoo [1987], Tables 2 and 3. For 4 variables the 5% critical value for the test is approximately -4.28 and for the ADF test is -4.0.

Sticky Price Monetary Model: The cointegration results of model (2) are shown in Table 4. The expected inflation differential is significant and has improved the explanatory power of all regressions but has made no difference to the cointegration tests.⁹ Both the DF and ADF tests strongly reject the existence of a cointegrating relationship between the variables for any exchange rate. These results are similar to those found by Baillie and Selover [1987] and McNown and Wallace [1989] for bilateral exchange rates against the US dollar, and undermine the findings obtained by past studies

Table 4^a
Cointegrating Regressions
Sticky Price Monetary Model: Equation (2)

Variables	US Dollar/ pound	D. Mark/ pound	Yen/ pound	F. Franc/ pound
Constant	-6.589 (1.49)	16.949 (5.19)	4.079 (15.52)	2.281 (4.49)
$(m^* - m)_t^s$	0.031 (0.24)	0.828 (10.97)	1.030 (5.59)	-0.008 (0.11)
$(y^* - y)_t$	-1.600 (1.94)	1.393 (2.62)	-0.594 (1.41)	
$(r^* - r)_t$	2.646 (3.49)	3.503 (4.28)	1.567 (1.70)	0.209 (0.37)
$(\pi^* - \pi)_t$	-3.219 (2.72)	2.603 (2.22)	4.998 (2.67)	3.223 (3.77)
\bar{R}^2	0.495	0.744	0.853	0.396
DF	-2.078	-1.647	-2.032	-2.397
[LM(4)]	[6.2]	[12.6]	[16.7]	[7.43]
ADF (k)	-2.330(2)	-2.508(4)	-1.789(3)	-2.676(1)
[LM(4)]	[3.9]	[3.1]	[5.1]	[4.94]

Notes: (a) For a description of the diagnostic statistics, see Table 3. For 5 variables, the 5% critical value for the DF test is approximately -4.67 and for the ADF test is -4.25 (Engle and Yoo [1987], Table 3).

9. The use of the long-term interest rate as a proxy for unobservable expected price inflation follows a similar approach by Meese and Rogoff [1983], Baillie and Selover [1987], and Frankel [1979].

using monetarist models.

Hooper-Morton Model: The new variable in equation (3) is the cumulated current account balance. The integration tests in Section IV show that this variable is at least I(2) in all countries, except for Japan, and therefore cannot be integrated with any currency. This consistent result for all countries undermines the findings by Hooper and Morton [1982], and Meese and Rogoff [1983], who estimated equation (3) using domestic and foreign cumulated current account balances. In an effort to obtain some evidence for the HM model, we decided to use the stocks of foreign assets held by

Table 5^a
Cointegrating Regressions
Hooper-Morton Model: Equation (3)

Variables	US Dollar/ pound	D. Mark/ pound	Yen/ pound	F. Franc/ pound
Constant	3.852 (1.22)	7.084 (22.30)	7.197 (19.00)	0.942 (1.48)
$(m^* - m)_t^s$	0.646 (10.10)	0.324 (2.53)	0.437 (3.17)	-0.250 (2.45)
$(y^* - y)_t$	-0.762 (1.47)	-0.637 (2.51)	-0.357 (1.28)	
$(r^* - r)_t$	0.627 (0.83)	-0.371 (0.64)	0.707 (1.07)	0.468 (0.87)
$(\pi^* - \pi)_t$	2.637 (2.92)	6.545 (5.78)	5.968 (4.81)	3.690 (4.52)
f^{*t}	-0.103 (6.81)	-0.183 (10.90)		-0.067 (3.18)
$(cca)^s$			-0.052 (9.29)	
\bar{R}^2	0.848	0.947	0.936	0.468
DF	-4.087	-3.210	-2.686	-2.353
[LM(4)]	[3.25]	[11.11]	[12.40]	[8.39]
ADF(k)	-3.286(4)	-3.378(3)	-3.332(3)	-2.76(1)
[LM(4)]	[2.95]	[1.33]	[1.44]	[3.95]

Notes: (a) For a description of diagnostic statistics, see table 3.

domestic residents, f and f^* , instead of (cca) and $(cca)^*$, provided that they are $I(1)$. In the case of Japan we used both f^* and $(cca)^*$. The tests, shown in Table 5, reject overwhelmingly the existence of a cointegrating relationship between the variables for any exchange rate.

Portfolio Balance Model: All variables included in the exchange rate equation (4) are $I(1)$, including the seasonally adjusted series, so we can

Table 6^a
Cointegrating Regressions
Portfolio Balance Model: Equation (4)

Variables	US Dollar/ pound	D. Mark/ pound	Yen/ pound	F. Franc/ pound
Constant	-7.850 (4.75)	0.151 (0.11)	0.070 (0.03)	-0.095 (0.09)
$(p^* - p)_t$	1.730 (6.46)	0.752 (3.84)	1.054 (4.12)	0.982 (8.58)
m_t^s	-0.209 (2.44)	-0.015 (0.23)	0.237 (3.47)	
$(m^* - m)_t^s$				-0.034 (0.30)
m_t^s	0.791 (5.62)	0.012 (0.10)	-0.312 (1.95)	
k_t^s	-0.221 (0.61)	-0.030 (1.17)	0.018 (0.61)	-0.033 (1.72)
b_t^s	-0.057 (1.59)	0.051 (2.83)	0.053 (2.69)	0.066 (3.26)
f_t^*		-0.048 (4.07)	-0.110 (4.73)	0.048 (1.57)
$(vnso)_t$	-3.955 (6.23)	0.573 (1.30)	1.233 (2.75)	2.123 (5.29)
\bar{R}^2	0.727	0.926	0.960	0.639
DF	-2.448	-4.402	-3.154	-4.232
[LM(4)]	[8.393]	[2.04]	[9.65]	[7.03]
ADF (k)	-3.41(1)	-4.30(1)	-4.21(1)	-4.754
[LM(4)]	[3.34]	[5.52]	[5.35]	[6.60]

Notes: (a) For a description of the diagnostic statistics, see Table 3.

Table 7^a
Johansen Maximum Likelihood Test for Cointegration (-2lnQ)

	Null	D-Mark/ Pound	Yen/ Pound	F-Franc/ Pound	5% Critical Value
A. Portfolio Balance Model(4)	$z = 0$	75.830	74.048	69.001	51.420
	$z \leq 1$	42.544	37.127	39.034	45.277
	$z \leq 2$	30.841	34.148	29.580	39.372
	$z \leq 3$	22.478	18.167	24.364	33.461
	$z \leq 4$	18.539	12.873	20.650	27.067
	$z \leq 5$	13.268	9.186	17.818	20.967
B. Modified UIP Model (7)	$z = 0$	40.756	42.886	45.722	39.372
	$z \leq 1$	28.436	30.316	29.351	33.461
	$z \leq 2$	16.311	22.425	17.394	27.067
	$z \leq 3$	12.043	13.694	15.252	20.967
	$z \leq 4$	9.843	10.103	8.021	14.069

Notes: (a) $-2\ln Q$ is the Likelihood Ratio test statistic that there are z number of cointegrating vectors (Johansen [1988]).

proceed to testing for cointegration.¹⁰ The results are shown in Table 6. Both the DF and ADF tests reject the existence of a cointegrating vector between the variable for the dollar/pound exchange rate. In the case of the other three exchange rates, there seems to be a conflict between the DF and ADF statistics. But in view of the LM autocorrelation results, the ADF statistic is the more appropriate test. This test strongly supports the existence of a cointegrating relationship for the franc/pound exchange rate. In the case of the D-mark and yen, the ADF statistic is on the borderline which provides tentative evidence in support of a cointegrating relationship. The Johansen tests reported in Table 7 also confirm the uniqueness of these

10. Unfortunately we have been unable to obtain data on K^* and B^* for USA, Germany, Japan and France, so these variables are not included in the estimation. One might argue that we should have used bilateral data for F and F^* , but the unavailability of such data has forced us to use total private foreign assets. A similar approach has been adopted by Branson *et al.* [1977] and Sarantis [1987]. It should also be noted that m^* for France is $I(2)$ and cannot be cointegrated with the exchange rate, so we have instead used relative money $(m^* - m)^*$ which is $I(1)$.

cointegrating vectors.

Modified Uncovered Interest Parity (Forward-Looking): Equation (5) suggests that it is the change in the exchange rate that it is related to the interest rate differential and the risk premium term (as well as to the expected inflation differential and relative prices in the case of the real version of (5)).¹¹ Therefore if the change in e is stationary, the interest rate differential and the other variables should also be stationary (*i.e.* $I(0)$) in order to satisfy Granger's requirement that the two sides of the equation should match up. Our statistical tests show that while e is $I(1)$ and so the change in e is stationary, none of the other variables is $I(0)$ – they are all $I(1)$. This statistical evidence on the integrability of the variables rejects the existence of an exchange rate equation consistent with the forward-looking UIP relationship for all bilateral exchange rates.¹²

Modified Uncovered Interest Parity (Backward-Looking): The estimates of equation (7) are shown in Table 8. Again the DF and ADF statistics strongly reject the existence of a cointegrating vector for the dollar/pound exchange rate. With respect to the other three exchange rates, the LM statistic for the DF-equation shows significant autocorrelation, so the appropriate test in these cases is the ADF test. The relevant statistics suggest that the variables cointegrate for all three exchange rates, with the stronger evidence for the yen. The uniqueness of these cointegrating vectors is strongly supported by the Johansen tests shown in Table 7. The likelihood ratio (LR) tests also reject the restriction of a unit coefficient on the interest rate differ-

11. We are grateful to K. Wallis and S. Hall for stressing this point to us in private correspondence.

12. Fisher *et al* [1990] have argued that interest rate differentials cannot drift apart without bound, and that over a longer period the interest rate differential is $I(0)$. All statistical results are sample specific and ours are not an exception. Nevertheless we have investigated this argument by computing the ADF tests for both the interest rate differential and expected inflation differential (proxied by the long term interest rate differential) over different sample periods: 1971Q2-1990Q3, 1973Q1-1990Q3, 1977Q1-1990Q3, and 1980Q1-1990Q3. The results show that the hypothesis of $I(0)$ is rejected for all interest rate differentials and irrespective of the sample period used, except in the case of Germany for the shorter period 1977Q1-1990Q3. These results reject the argument by Fisher *et al*. The differences may of course be due to different measures of interest rates used.

Table 8^a
Cointegrating Regressions
Modified Uncovered Interest Parity: Equation (7)

Variables	US Dollar/ pound	D. Mark/ pound	Yen/ pound	F. Franc/ pound
Constant	0.689 (16.80)	1.427 (48.90)	5.973 (150.00)	2.333 (256.10)
$(r^* - r)_t$	1.468 (2.01)	0.195 (0.37)	0.482 (1.22)	-0.242 (0.66)
$(\pi^* - \pi)_t$	-0.491 (0.52)	2.703 (4.29)	1.389 (1.85)	4.337 (9.29)
$(p^* - p)_t$	0.085 (0.71)	0.627 (19.00)	0.733 (22.40)	0.647 (7.53)
$(CA/NY)_t^{*s}$	9.330 (5.37)	-0.734 (4.60)	-1.781 (11.80)	-0.712 (1.63)
$(CA/NY)_t$	-2.432 (3.46)	-0.480 (0.88)	1.838 (5.56)	-0.647 (1.53)
\bar{R}^2	0.689	0.927	0.975	0.714
DF	-2.801	-2.910	-4.608	-3.384
[LM(4)]	[4.04]	[14.03]	[14.92]	[8.23]
ADF (k)	-2.800(4)	-4.30(4)	-6.382(4)	-4.381(4)
[LM(4)]	[2.78]	[6.29]	[2.54]	[7.46]

Notes: (a) For a description of diagnostic statistics, see Table 3. $(CA/NY)^*$ for the USA is unfiltered.

ential predicted by the UIP hypothesis; the LR(1) statistics for the D-mark/pound, yen/pound and F-franc/pound rates are 11.38, 7.65 and 5.98 respectively. Our findings contrast with those by Fisher *et al.* [1990] for sterling's real effective exchange rate. It is also interesting to notice that strong evidence in favor of using backward-looking exchange rate expectations among foreign exchange dealers in the City of London is provided in a recent survey by Allen and Taylor [1989]. Similar evidence was found by Frankel and Froot [1987] for USA foreign exchange dealers.

V. Short-Run Dynamic Equations

In this section we develop short-run dynamic forms of the exchange rates based on the cointegrating equations obtained in the previous section. Following the two-step procedure of Engle and Granger [1987], the lagged residuals RES_{t-1} from the cointegrating or long-run relationship are entered into the error correction model (ECM) and should display a negative sign. In specifying the dynamic model we have used the general to specific methodology, starting with lags of up to four quarters of the change in the variables included in the respective cointegrating equation and nested down to a parsimonious representation of the data.¹³ The error correction models were estimated with the OLS method. These regressions [(8)-(13)] include the contemporaneous values of some explanatory variables. Unless these variables are independent of the disturbance term, the OLS estimates will be inconsistent. We have therefore tested for exogeneity using both the Granger-causality test and the Wu-Hausman (Wu [1973]) exogeneity test. The statistics reported in Table 9 show that none of the contemporaneous variables is Granger caused by the exchange rate. Furthermore, the exogeneity hypothesis for all variables is strongly supported by the Wu-Hausman statistics for all models, which suggests that the OLS estimates of the dynamic equations are consistent. The error correction models were estimated over the period 1973Q2 - 1989Q3, with the four observations for 1989Q4 - 1990Q3 used for ex post forecasting tests.

A. Dollar/Pound

According to the Granger representation theorem, for a valid ECM to exist the set of variables must cointegrate. In the case of the dollar/pound exchange rate we have failed to find statistical evidence in support of a cointegrating or long-run equilibrium relationship consistent with any theoretical exchange rate model. Consequently, a valid ECM representation for the dollar/pound exchange rate cannot exist.

13. For a good exposition of the 'general to specific' methodology and an explanation of the diagnostic statistics reported with regressions (8)-(13), see Cuthbertson *et al.* [1992]. The same reference also provides an excellent analysis of the cointegration and error correction formulations.

Table 9^a
Exogeneity Tests (1973Q2-1990Q3)

A. Modified UIP Model						
A1. D-Mark/ Pound		$\Delta(\pi^* - \pi)$	$\Delta(CA/NY)^s$	$\Delta(CA/NY)$	All Variables	
	Granger-Causality	0.421	2.155	0.836	-	
	Wu-Hausman	-	-	-	0.522	
A2. Yen/Pound		$\Delta(\pi^* - \pi)$	$\Delta(CA/NY)^s$	$\Delta(CA/NY)$	All Variables	
	Granger-Causality	1.708	0.523	1.202	-	
	Wu-Hausman	-	-	-	0.183	
A3. F-Franc/ Pound		$\Delta(r^* - r)$	$\Delta(\pi^* - \pi)$	$\Delta(CA/NY)^s$	All Variables	
	Granger-Causality	1.023	0.319	0.796	-	
	Wu-Hausman	-	-	-	0.344	
B. Portfolio Balance Model						
B1. D-Mark/ Pound		$\Delta(p^* - p)$	Δm^s	Δb^s	Δf^*	All Variables
	Granger-Causality	2.472	0.388	1.720	2.057	-
	Wu-Hausman	-	-	-	-	1.052
B2. Yen/Pound		$\Delta(p^* - p)$	Δm^s	Δk^s	Δf^*	All Variables
	Granger-Causality	1.388	0.623	0.447	1.610	-
	Wu-Hausman	-	-	-	-	0.633
B3. F-Franc/ Pound		Δk^s	Δb^s	Δf^*	All Variables	
	Granger-Causality	1.061	1.615	0.497	-	
	Wu-Hausman	-	-	-	0.149	

Notes: (a) The 5% critical value for the Granger-causality statistic is $F(4,61)=2.53$. The 5% critical values for the Wu-Hausman statistics are approximately as follows: A1: $F(3,57)=2.82$; A2: $F(3,57)=2.82$; A3: $F(3,56)=2.82$; B1: $F(4,55)=2.59$; B2: $F(4,52)=2.57$; B3: $F(3,59)=2.75$.

B. D-Mark/Pound

Modified Uncovered Interest Parity: Using the residuals (RES) from the corresponding cointegrating regression in Table 8 and following the methodology outlined above has produced the following dynamic model (*t*-values in parentheses).

$$\Delta e_t = 0.004 - 1.009 \Delta(r^* - r)_{t-1} + 2.925 \Delta(\pi^* - \pi)_t + 0.514 \Delta(p^* - p)_{t-4}$$

(0.60) (3.70) (5.73) (1.83)

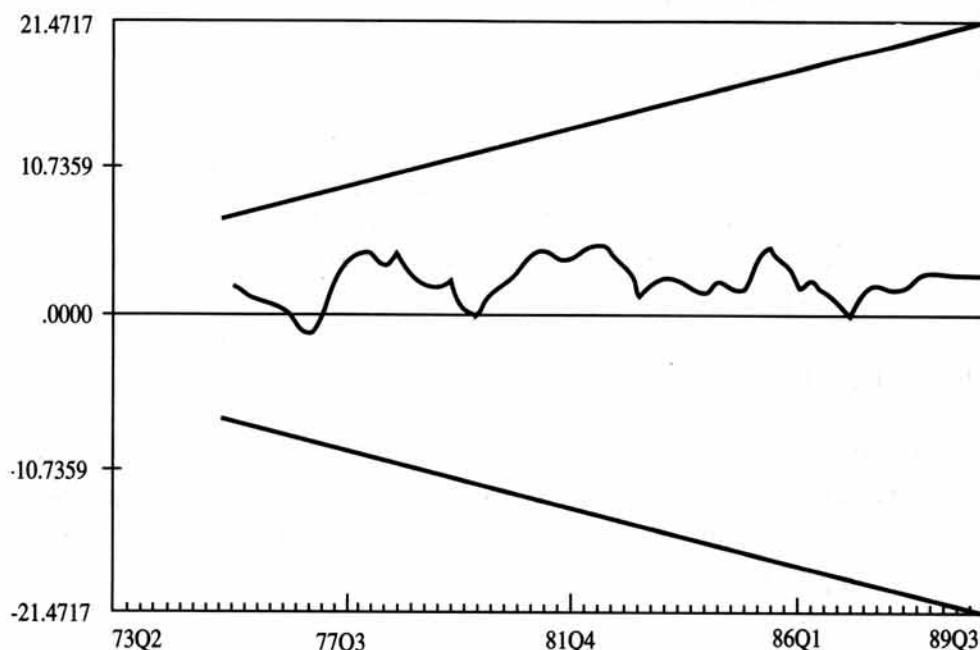
$$\begin{aligned}
& -0.420 \Delta(CA/NY)_t^s + 0.508 \Delta(CA/NY)_t + 0.307 \Delta e_{t-1} \\
& \quad (1.29) \qquad \qquad (1.79) \qquad \qquad (3.07) \\
& +0.168 \Delta e_{t-3} - 0.247 RES_{t-1} \\
& \quad (1.96) \qquad \qquad (3.55)
\end{aligned} \tag{8}$$

$$\begin{aligned}
\bar{R}_2 &= 0.54, \quad \sigma = 0.031, \quad DW = 1.91, \quad LM_a(4) = 6.05 \\
LM_f(1) &= 0.77, \quad LM_n(2) = 0.98, \quad LM_h(1) = 6.68 \\
ARCH(1) &= 0.30, \quad ARCH(4) = 0.22, \quad LM_p(4) = 6.64 \\
STAB[79Q2] &= 1.97(9,52), \quad STAB[84Q4] = 1.47(9,52)
\end{aligned}$$

For definitions of the diagnostic test statistics see Appendix I.

The diagnostic tests show no sign of misspecification. The residuals of the model are uncorrelated, homoscedastic and normally distributed; the functional form is supported even at a 1% level of significance; and the post sample predictive performance is satisfactory. Given the notorious structural instability of exchange rate equations we decided to investigate the parameter stability of all dynamic equations further than relying on the post

Figure 1
Plot of Cumulative Sum of Recursive Residuals: Model (8)



The Straight lines represent critical bounds at 5% significance level

sample prediction test only. We conducted structural stability tests by splitting the sample both in 1979Q2 which was the beginning of the European Monetary System, and in 1984Q4 when sterling reached historically its lowest value. Both statistics are insignificant, thus suggesting parameter stability. This result is also supported by the recursive residuals which are reasonably random (the CUSUM of recursive residuals remains within the 5% significance lines, Figure 1).

All regression coefficients are correctly signed, including the error correction term (RES_{t-1}) which is highly significant at both 5% and 1% levels of significance. With regards to the other variables, relative interest rates and expected inflation differentials exert the strongest influence.

Portfolio Balance Model: Entering the residuals from the corresponding cointegrating regression in Table 6 into the ECM formulations has produced the model

$$\begin{aligned} \Delta e_t = & 0.020 + 0.449\Delta(p^* - p)_t - 0.215\Delta m_{t-2}^s - 0.263\Delta m_t^s \\ & (1.44) \quad (1.20) \quad (1.69) \quad (2.01) \\ & - 0.059\Delta k_{t-4}^s + 0.049\Delta b_t^s + 0.081\Delta b_{t-4}^s - 0.017\Delta f_t^* \\ & (1.85) \quad (1.60) \quad (2.81) \quad (2.13) \\ & + 0.269\Delta e_{t-1} - 0.417\Delta RES_{t-1} \end{aligned} \quad (9)$$

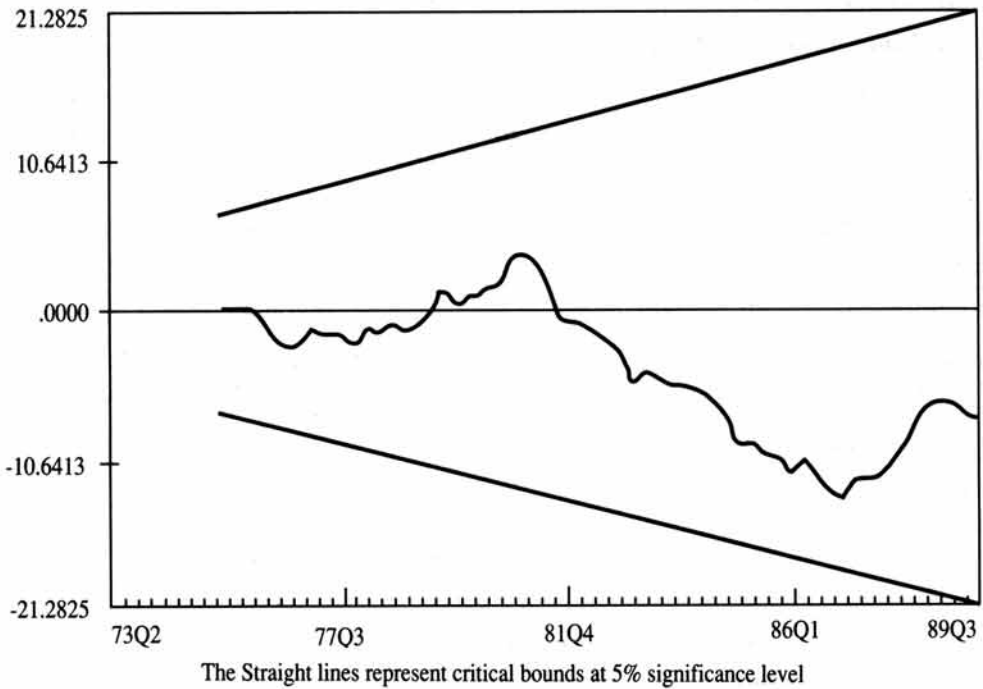
(2.15) (4.01)

$$\begin{aligned} \bar{R}_2 &= 0.28, \quad \sigma = 0.039, \quad DW = 1.86, \quad LM_a(4) = 2.01 \\ LM_f(1) &= 4.82, \quad LM_n(2) = 0.23, \quad LM_h(1) = 0.35 \\ ARCH(1) &= 1.85, \quad ARCH(4) = 5.99, \quad LM_p(4) = 5.87 \\ STAB[79Q2] &= 1.10(10,50), \quad STAB[84Q4] = 0.89(10,50) \end{aligned}$$

The model passes the various diagnostic tests of its error terms and stability. The latter is supported not only by the prediction and structural stability tests, but also by the CUSUM of recursive residuals (shown in Figure 2). On the other hand, the $LM_f(1)$ statistic for functional form fails a 5% significance test (though it passes a 1% significance test). The error-correction term has a negative and highly significant coefficient. The other variables have the anticipated signs, except German money. North sea oil had a t-value well below unity in all experiments and was therefore dropped.

How does this model compare with the modified UIP equation (8)? Both

Figure 2
Plot of Cumulative Sum of Recursive Residuals: Model (9)



models appear to be satisfactory on the basis of the diagnostic tests, but the explanatory power of the BP model is relatively poor in comparison to that of model (8). A more formal comparison is provided by the non-nested tests shown in Table 10. Godfrey and Pesaran [1983], who provide an explanation and useful comparison of these non-nested tests, suggest that the W- and JA-tests perform better in small samples. Both these tests strongly accept model M_1 (8) and reject M_2 (9). The encompassing test is less clear. The statistic for M_1 vs M_2 fails a 5% test but passes a 1% test, whereas the statistic for M_2 vs M_1 is strongly rejected at both 5% and 1% significance levels. Combined with the Akaike and Schwarz Information Criteria, the overall evidence obtained by the non-nested tests favors the dynamic MUIP model.

C. Yen/Pound

Modified Uncovered Interest Parity: Using the residuals from the co-integrating regression for the yen/pound rate in Table 8 as the error correction term, the general to specific methodology has produced the following

Table 10
Non-Nested Tests for Dynamic Models (8) and (9)

Test Statistic	M_1 against M_2	M_2 against M_1
N-Test	-3.94	-13.61
NT-Test	-2.44	-8.69
W-Test	-2.22	-6.27
J-Test	3.64	7.60
JA-Test	1.49	4.72
Encompassing	F(8,49) = 2.41	F(7,49) = 8.02
Akaike's Information Criterion of M_1 vs. M_2 = 15.23 (favours M_1).		
Schwarz's Bayesian Information Criterion of M_1 vs. M_2 = 16.32 (favours M_1)		
M_1 = model (8), M_2 = model (9)		

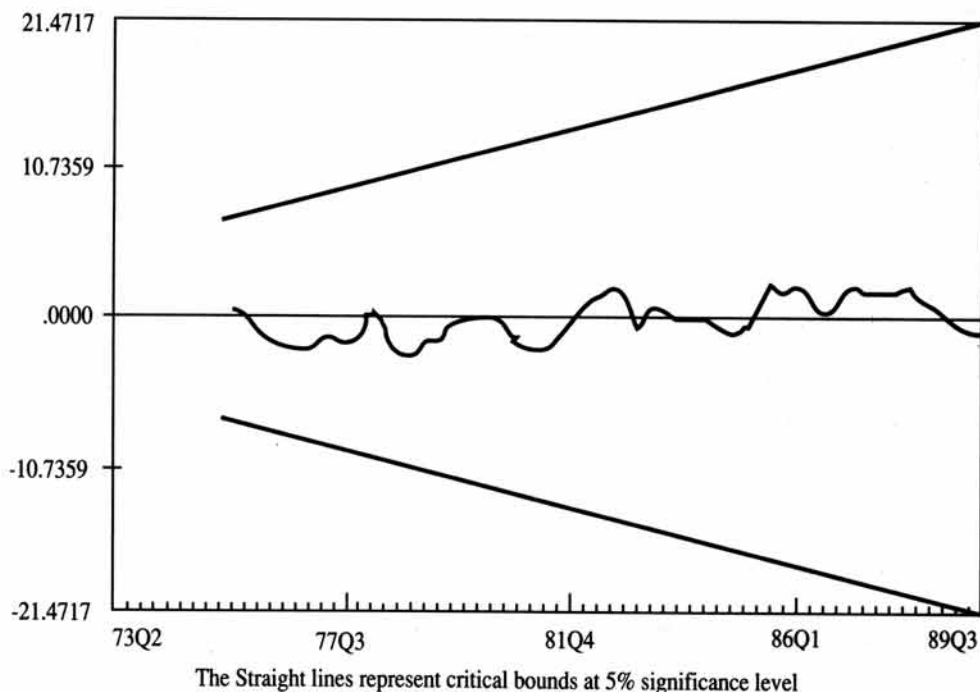
dynamic model:

$$\begin{aligned}
 \Delta e_t = & \underset{(1.68)}{-0.008} - \underset{(2.26)}{0.828\Delta(r^* - r)_{t-1}} + \underset{(3.80)}{2.279\Delta(\pi^* - \pi)_t} \\
 & - \underset{(2.77)}{1.032\Delta(CA/NY)_t^s} + \underset{(2.56)}{1.014(CA/NY)_t} - \underset{(1.99)}{0.693\Delta(CA/NY)_{t-1}} \\
 & + \underset{(1.09)}{0.405\Delta(CA/NY)_{t-2}} + \underset{(4.22)}{0.477\Delta e_{t-1}} - \underset{(4.74)}{0.449 RES_{t-1}} \quad (10)
 \end{aligned}$$

$$\begin{aligned}
 \bar{R}_2 &= 0.50, \quad \sigma = 0.036, \quad DW = 2.05, \quad LM_a(4) = 6.89 \\
 LM_f(1) &= 0.67, \quad LM_n(2) = 1.40, \quad LM_h(1) = 0.44 \\
 ARCH(1) &= 4.47, \quad ARCH(4) = 6.33, \quad LM_p(4) = 6.01 \\
 STAB[79Q2] &= 0.62(9,52), \quad STAB[84Q4] = 1.03(9,52)
 \end{aligned}$$

The model passes a wide range of diagnostic tests of its error terms, functional form and parameter stability. The latter is particularly remarkable irrespective of whether we split the sample in 1979 or 1984. This finding is also supported by the near to zero randomness of the recursive residuals, as shown by the CUSUM test in Figure 3. The error-correction term (RES_{t-1}) has the anticipated negative effect and is highly significant at the 1% level. All other variables are correctly signed and significant. The only exception is the relative price which had consistently a t -value below unity in all esti-

Figure 3
Plot of Cumulative Sum of Recursive Residuals: Model (10)



mations and has been dropped.

Portfolio Balance Model: Once again we proceed to a dynamic model which is based on the cointegrating regression for the yen/pound rate shown in Table 6. The final model is:

$$\begin{aligned}
 \Delta e_t = & 0.011 + 1.224\Delta(p^* - p)_t - 0.286\Delta m_t^s + 0.453\Delta m_{t-1}^s + 0.040\Delta k_t^s \\
 & (1.05) \quad (2.81) \quad (1.83) \quad (2.73) \quad (1.31) \\
 & - 0.065\Delta k_{t-4}^s - 0.0426\Delta b_{t-3}^s - 0.104\Delta f_t^* - 0.081\Delta f_{t-2}^* + 0.048\Delta f_{t-4}^* \\
 & (2.35) \quad (1.85) \quad (2.75) \quad (2.25) \quad (1.40) \\
 & + 1.542\Delta(vnso)_{t-2} + 0.331\Delta e_{t-1} - 0.387 RES_{t-1} \quad (11) \\
 & (2.58) \quad (2.84) \quad (3.64)
 \end{aligned}$$

$$\bar{R}_2 = 0.44, \quad \sigma = 0.039, \quad DW = 1.86, \quad LM_a(4) = 1.89$$

$$LM_f(1) = 0.01, \quad LM_n(2) = 1.79, \quad LM_h(1) = 0.01,$$

$$ARCH(1) = 5.48, \quad ARCH(4) = 6.84, \quad LM_p(4) = 6.71$$

$$STAB[79(2)] = 0.89(13,44), \quad STAB[84Q4] = 1.30(13,44)$$

All the diagnostic tests are satisfactory, thus suggesting acceptance of the parsimonious model. The CUSUM statistic obtained from the recursive estimation (shown in Figure 4) indicates parameter stability, though the CUSUM test is not as close to the zero line as that for model (10). The coefficients on all variables have the anticipated signs and are significant, except those on money. The money supply for Japan was consistently insignificant (with a *t*-value below unity) irrespective of the lag structure and was therefore dropped. The UK money supply is significant but wrongly signed. It is interesting to notice here the strong influence of North Sea oil.

A formal comparison of the dynamic models (10) and (11) is carried out with the non-nested tests shown in Table 11. The W- and JA-tests seem to suggest acceptance of both models so they cannot discriminate between the two. On the other hand, the encompassing test clearly accepts model M_1 (10) and rejects M_2 (11). The same result is obtained by the Akaike and Schwarz Information Criteria, so the modified UIP model (10) appears to outperform the dynamic PB model (11).

Figure 4
Plot of Cumulative Sum of Recursive Residuals: Model (11)

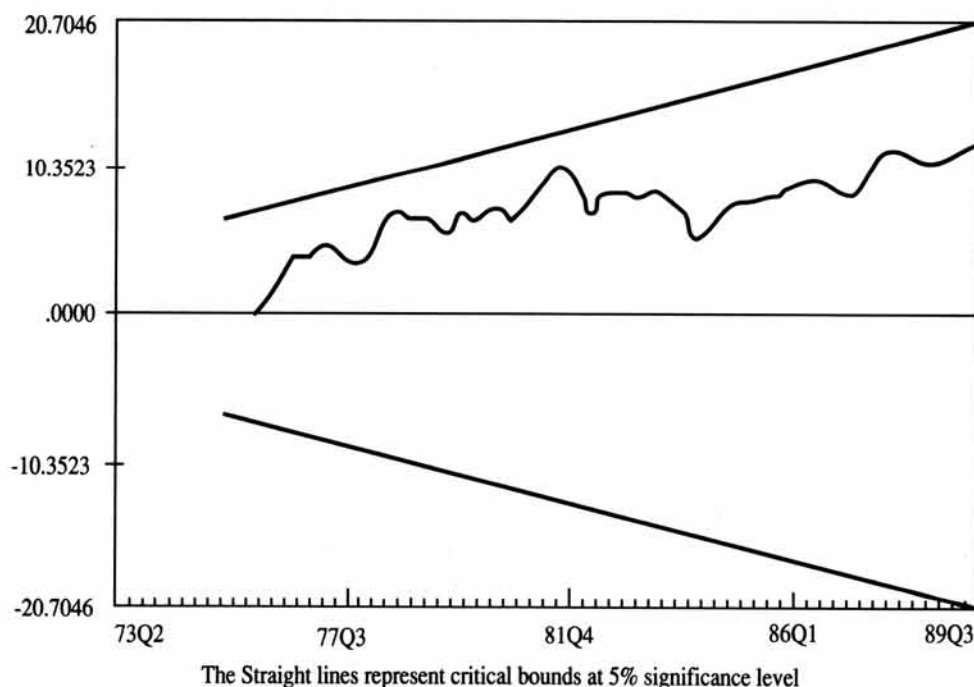


Table 11
Non-Nested Tests for Dynamic Models (10) and (11)

Test Statistic	M_1 against M_2	M_2 against M_1
N-Test	-6.65	-7.06
NT-Test	-3.60	-4.52
W-Test	-3.07	-3.72
J-Test	4.55	4.61
JA-Test	3.48	3.61
Encompassing	$F(11,46) = 1.72$	$F(7,46) = 3.11$
Akaike's Information Criterion of M_1 vs. $M_2 = 5.40$ (favours M_1).		
Schwarz's Bayesian Information Criterion of M_1 vs. $M_2 = 9.78$ (favours M_1)		
M_1 = model (10), M_2 = model (11)		

D. F-Franc/Pound

Modified Uncovered Interest Parity: Using the lagged residuals from the cointegrating regression for the franc/pound rate shown in Table 8 as an explanatory variable in the error-correction model has yielded the dynamic equation.

$$\begin{aligned}
 \Delta e_t = & -0.001 + 0.526\Delta(r^* - r)_t - 0.960\Delta(r^* - r)_{t-1} + 2.623\Delta(\pi^* - \pi)_t \\
 & (0.14) \quad (1.55) \quad (3.31) \quad (4.37) \\
 & + 0.443\Delta(p^* - p)_{t-1} - 0.489\Delta(p^* - p)_{t-3} + 0.316\Delta(p^* - p)_{t-4} \\
 & (2.01) \quad (2.18) \quad (1.50) \\
 & - 0.696\Delta(CA/NY)_t^* - 0.922_t(CA/NY)_{t-1} \\
 & (1.95) \quad (3.08) \\
 & + 0.282\Delta e_{t-1} - 0.230 RES_{t-1} \quad (12) \\
 & (2.83) \quad (3.33)
 \end{aligned}$$

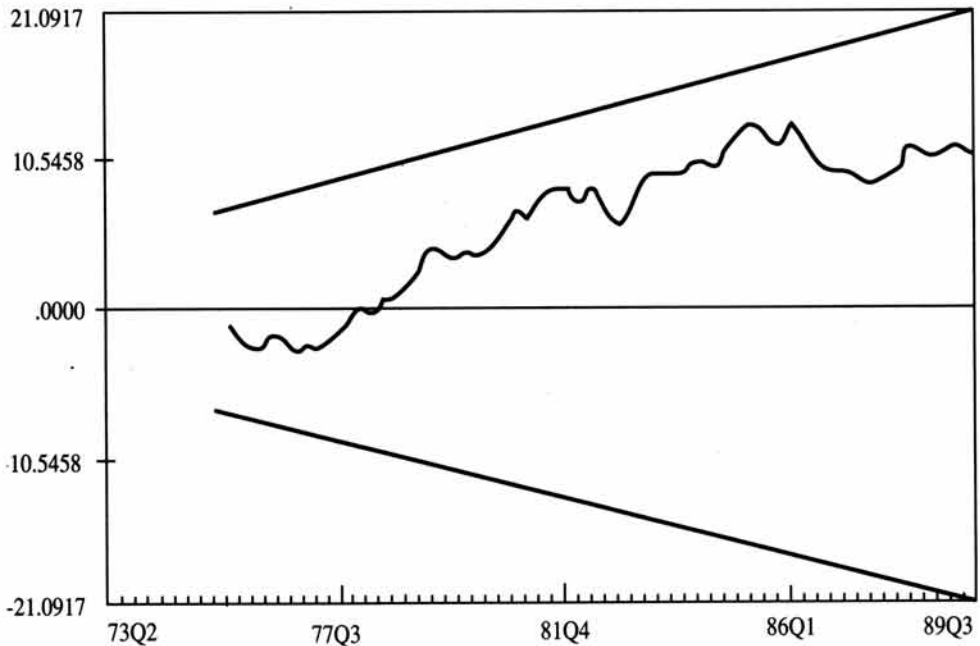
$$\bar{R}_2 = 0.53, \quad \sigma = 0.031, \quad DW = 1.78, \quad LM_a(4) = 2.75$$

$$LM_r(1) = 0.99, \quad LM_\pi(2) = 1.24, \quad LM_h(1) = 0.74,$$

$$ARCH(1) = 3.63, \quad ARCH(4) = 6.20, \quad LM_p(4) = 5.80$$

$$STAB[79Q2] = 0.51(11,48), \quad STAB[84Q4] = 1.31(11,48)$$

Figure 5
Plot of Cumulative Sum of Recursive Residuals: Model (8)



The Straight lines represent critical bounds at 5% significance level

There are no signs of misspecification, as the model passes a wide range of diagnostic statistics of its functional form, error term, ex post forecasting and stability. Irrespective of whether we split the sample in 1979 or 1984, the tests indicate remarkable structural stability. This is also considered by the CUSUM of the recursive residuals (shown in Figure 5), which is within its 5% confidence lines. The lagged residuals $(RES)_{t-1}$, have the anticipated negative sign and are highly significant at the 1% level. With respect to parameters on other variables, they are all correctly signed and significant, with the exception of the UK current account to GDP ratio which displays the wrong sign.

Portfolio Balance Model: The parsimonious dynamic model built on the basis of the cointegrating regression for the franc/pound rate shown in Table 6 has the form:

$$\Delta e_t = -0.011 - 0.059\Delta k_t^s + 0.085\Delta b_t^s + 0.072\Delta b_{t-4}^s$$

(1.86)
(2.04)
(2.93)
(2.70)

$$-0.045\Delta f_t^* + 0.354\Delta e_{t-1} - 0.428 RES_{t-1} \quad (13)$$

(1.70) (3.02) (4.70)

$$\bar{R}_2 = 0.29, \quad \sigma = 0.038, \quad DW = 1.91, \quad LM_a(4) = 1.58$$

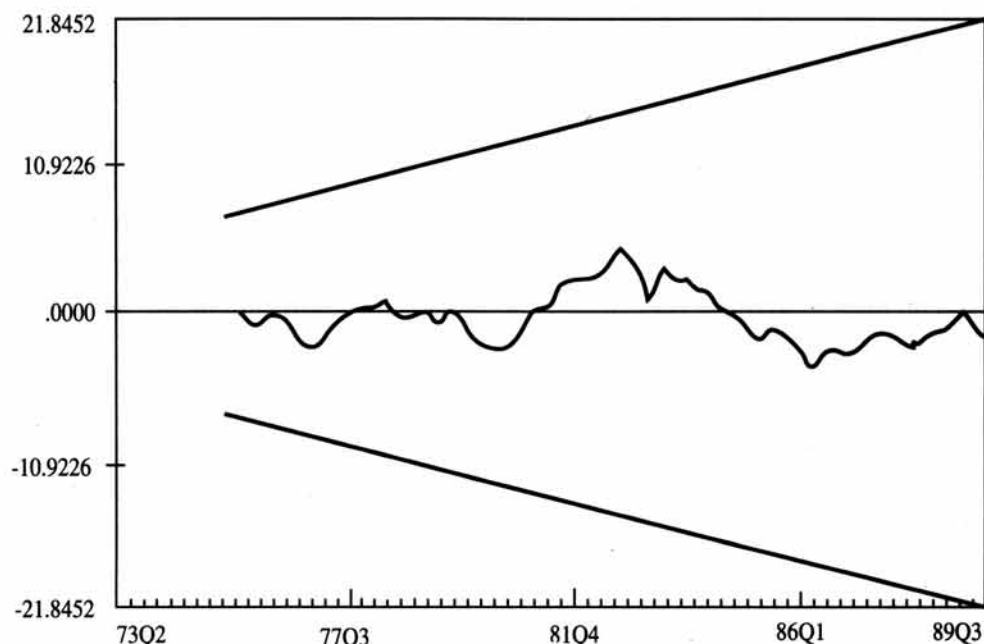
$$LM_f(1) = 3.10, \quad LM_n(2) = 0.45, \quad LM_h(1) = 2.03,$$

$$ARCH(1) = 0.20, \quad ARCH(4) = 0.29, \quad LM_p(4) = 5.84,$$

$$STAB[79Q2] = 1.20(7,56), \quad STAB[84Q4] = 0.49(7,56)$$

The model passes all the diagnostic tests and the CUSUM of the recursive residuals (shown in Figure 6) is close to zero. Hence the specification of the model appears to be satisfactory, but on the other hand many variables were persistently insignificant (with t -values below unity irrespective of the lag structure), so they were dropped from the final equation, and the explanatory power of the model is poor relative to that of model (2). It is interesting to notice that neither money stock, nor relative prices or North Sea oil exert any influence on the short-run behavior of the franc/pound exchange rate.

Figure 6
Plot of Cumulative Sum of Recursive Residuals: Model (13)



The Straight lines represent critical bounds at 5% significance level

A more formal comparison of models (12) and (13) is provided by the non-nested tests shown in Table 12. The encompassing test suggests that neither model can account for the salient features of the other model. On the other hand the W- and JA-tests accept the M_1 model (12) but reject the M_2 model (13). Furthermore, both the Akaike and Schwarz Information Criteria clearly favor the dynamic MUIP model.

VI. Conclusions

In this paper we have applied the cointegration-error correction techniques to model four bilateral sterling exchange rates. In each case we have applied the main theories of exchange rate determination and provide an econometric evaluation of the rival models. Our investigation of the integrability of time series shows the presence of stochastic seasonality in some variables, particularly in money supply, current account, trade balance, bonds and equity assets. In these cases we have removed the seasonal unit roots through appropriate seasonal filters, for otherwise the estimates of the cointegrating relationships would have been inconsistent. The cointegration results appear to provide dismal evidence for the monetary approach to exchange rate determination. We have been unable to find any statistical

Table 12
Non-Nested Tests for Dynamic Models (12) and (13)

Test Statistic	M_1 against M_2	M_2 against M_1
N-Test	-4.51	-13.58
NT-Test	-3.09	-7.68
W-Test	-2.08	-5.61
J-Test	4.33	8.89
JA-Test	1.08	2.62
Encompassing	$F(5,50) = 4.99$	$F(9,50) = 8.01$
Akaike's Information Criterion of M_1 vs. $M_2 = 12.11$ (favours M_1).		
Schwarz's Bayesian Information Criterion of M_1 vs. $M_2 = 7.73$ (favours M_1).		
M_1 = model (12), M_2 = model (13)		

evidence in support of a long-run relationship consistent with the flexible price monetary model, or the sticky price monetary model, or the Hooper-Morton model. This evidence undermines the validity of monetarist models used in forecasting exchange rates. The cointegration tests have uncovered unique long-run relationships consistent with the portfolio balance model and a modified uncovered interest parity relationship with backward-looking expectations for three exchange rates: D-mark/pound, yen/pound and F-franc/pound. In the case of the dollar/pound rate we have been unable to find a set of cointegrating variables consistent with any model.

On the basis of the cointegrating regressions which support long-run relationships we have built short-run dynamic models for the three sterling exchange rates which pass a wide range of diagnostic tests. In particular the structural stability of these models is remarkable. The error-correction term in all dynamic models is negative and highly significant. The coefficients on other variables are broadly correctly signed and significant, especially in the cases of the MUIP models. The estimates for the dynamic MUIP models show that interest rate and expected inflation differentials exert the strongest influence. Relative prices are important for the mark/pound and franc/pound rates, but not for the yen/pound rate. This contrasts with their highly significant long-term effects on all exchange rates. There is also evidence that risk premia (measured by the current account to GDP ratio) play a significant role (both in the short and long term) in exchange rate determination, particularly in the case of the yen/pound rate.

In the case of the dynamic portfolio balance models, foreign assets, bond and equity holdings tend to exert the strongest influence on all three exchange rates. North Sea oil has a strong short term influence only on the yen/pound rate, while in the long term it affects significantly both the yen and franc, and less so the mark. The estimates of the money supply effects were rather disappointing with most of them wrongly signed and/or insignificant.

Although the dynamic portfolio balance and MUIP models for each exchange rate were satisfactory by a broad range of diagnostic tests, in terms of goodness of fit and regression coefficients (sign and significance) the portfolio balance models appear to be less satisfactory. A formal compar-

ison of the rival models with non-nested tests tends to favor the dynamic MUIP model with backward-looking expectations for the D-mark/pound, yen/pound and F-franc/pound rates.

Appendix I

Diagnostic Statistics

\bar{R}^2	Coefficient of multiple determination adjusted for degrees of freedom
σ	Standard error of the regression
DW	Durbin-Watson statistic for 1 st order residual autocorrelation
$LM_a(4)$	Lagrange Multiplier test for 4th order residual autocorrelation; $\chi^2(4)$
$LM_f(1)$	Ramsey's RESET test of functional form; $\chi^2(1)$
$LM_n(2)$	The Jarque-Bera test for normality of residuals, based on a test of skewness and kurtosis of residuals; $\chi^2(2)$
$LM_h(1)$	Lagrange Multiplier test for heteroscedasticity, based on the regression of squared residuals on squared fitted values; $\chi^2(1)$
$ARCH(v)$	The Engle test for autoregressive conditional heteroscedasticity; $\chi^2(v)$
$LM_p(4)$	Chow's test for post-sample predictive failure; $\chi^2(4)$
$STAB(v_1, v_2)$	Chow's F-test for structural stability based on sample splits in 1979Q2 and 1984Q4 respectively
CUSUM	Cumulative sum of recursive residuals
N-Test	The Cox non-nested test
NT-Test	The adjusted Cox test
W-Test	Wald-type non-nested test; $\chi^2(1)$
J-Test	Davidson-Mackinnon non-nested test; $t(1)$ (valid asymptotically)
JA-Test	Fisher-McAleer non-nested test; $t(1)$ (small sample test)
Encompassing Mizon-Richard encompassing test; $F(v_1, v_2)$	

All regression estimates and diagnostic statistics were derived using Microfit.

Appendix II

Measurement of Variables and Data Sources

All variables are seasonally unadjusted, except y and y^* .

- $e =$ (log) Foreign currency/pound sterling spot exchange rate, average of daily telegraphic transfer rates in London – *Financial Statistics*.
- $m =$ (log) UK money stock (M_1) in pounds¹⁴ – data supplied by the Economics Division of the Bank of England.
- $m^* =$ (log) Foreign money stock (M_1) in foreign currency¹⁴ (i.e. for the U.S.A., West Germany, Japan and France) – *OECD Main Economic Indicators*.
- $y =$ (log) UK real gross domestic product in pounds – *OECD Main Economic Indicators*.
- $y^* =$ (log) Foreign real gross domestic product in foreign currency – *OECD Main Economic Indicators*.
- $r^{15} =$ (log) UK interest rate on 3 month inter-bank loans, monthly averages – *Financial Statistics*.
- $r^* =$ (log) Foreign money market interest rate (Federal Funds interest rate for the U.S.A), monthly averages – *OECD Main Economic Indicators and IMF International Financial Statistics*.
- $\pi =$ (log) UK long-term interest rate (yield of long-term government bonds), monthly averages – *OECD Main Economic Indicators*.
- $\pi^* =$ (log) Foreign long-term interest rate (U.S.A: yield of government composite bonds > 10 years; Germany and Japan: yield of

14. The use of M_1 as a measure for stock money is based partly on supportive evidence by previous studies (Magee and Rogoff [1983], Sarantis [1987], Branson *et al.* [1977], Hooper and Morton [1982]) and partly on the need to maintain data consistency across countries. It should be noted that while broad money M_2 (the main alternative to M_1) is available for U.S.A., Germany, Japan and France, the UK does not publish data on this measure of money. Instead the UK uses M_3 and lately M_4 as measures for broad money neither of which are available for the other countries.

15. This variable is computed as follows. If R_t denotes the interest rate in per cent per period, then $r_t = \ln(1 + R_t/100)$. The variables r^* , π and π^* were also computed in the same way.

long-term government bonds; France: yield of long-term bonds guaranteed by government), monthly averages – *OECD Main Economic Indicators*.

p = (log) UK producer price index – *OECD Main Economic Indicators*.

p^* = (log) Foreign producer price index (GDP implicit price deflator for France) – *OECD Main Economic Indicators*.

NY = UK gross domestic product at current prices (in pounds) – *OECD Main Economic Indicators*.

NY^* = Foreign gross domestic product at current prices (in foreign currency) – *OECD Main Economic Indicators*.

CA = UK current account balance in pounds – *OECD Main Economic Indicators*.

$(CA)^*$ = Foreign current account balance in foreign currency (data for Japan were in US dollars and were converted in yen using the yen/dollar exchange rate) – *OECD Main Economic Indicators*.

$(cca)^{16}$ = (log) UK cumulated current account balance on benchmark observation for 1971Q1.

$(cca)^*$ = (log) Foreign cumulated current amount balance on benchmark observations for 1971Q1.

f = (log) Foreign asset stock held by the UK private sector, F , computed as follows:

$$F_t = F_{t-1} + [CA_t - \Delta(RESERV)_t]$$

where $\Delta(RESERV)$ is the change in non-gold foreign reserves (these were in US dollars and were converted into pounds using the dollar/pound rate). The benchmark was 1971Q1. For an explanation and use of this formula see Branson *et al.* [1977]. Data for RESERV were obtained from *IMF International Financial Statistics*.

f^* = (log) Foreign asset stock held by the private sector of the foreign country, F^* , computed by the formula:

16. It should be pointed out that because some observations for (CCA) are negative, the series was scaled by a constant so that we could take its logarithm. This does not affect the elasticities. The same procedure was adopted for $(cca)^*$, f and f^* .

$$F_t^* = F_{t-1}^* + [CA_t^* - \Delta(RESERV)_t^*]$$

where the change in non-gold foreign reserves was converted from U.S. dollars to the currency of the respective country using the corresponding exchange rate between the U.S. dollar and that country.

- k = (log) UK personal sector holdings of equity assets – data provided by the Centre for Economic Forecasting of the London Business School.
- b = (log) UK personal sector holdings of public sector debt - data provided by the Centre for Economic Forecasting of the London Business School.
- $vnso$ = Value of North Sea oil production in pounds - data provided by the Centre for Economic Forecasting of the London Business School - expressed as a percentage of UK nominal GDP.

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