

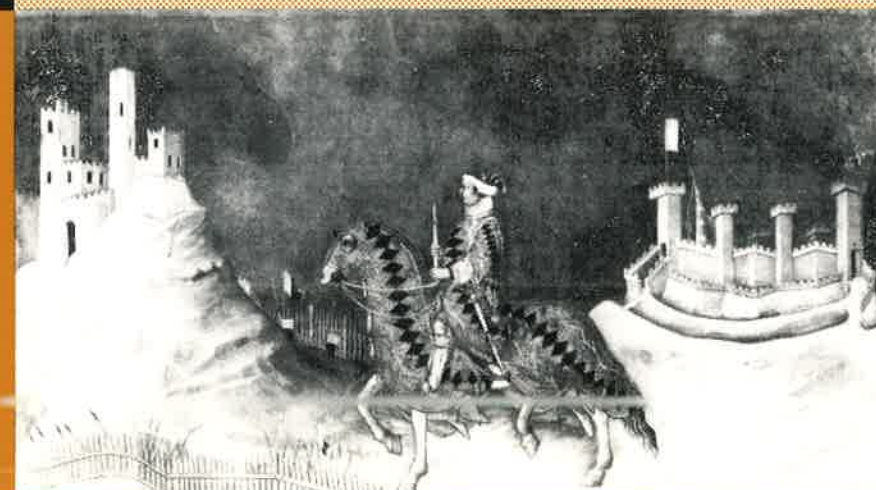
UNIVERSITA' DEGLI STUDI DI SIENA



QUADERNI DEL DIPARTIMENTO
DI ECONOMIA POLITICA

GIULIO CIFARELLI

**FUNDAMENTALS, REGIME SHIFTS
AND DOLLAR BEHAVIOUR IN THE 1980s.
AN EMPIRICAL ANALYSIS**



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Giulio Cifarelli

**FUNDAMENTALS, REGIME SHIFTS AN DOLLAR
BEHAVIOUR IN THE 19805
AN EMPIRICAL ANALYSIS**



Siena, ottobre 1992

The dollar behaviour in the 1980s has been extensively investigated. The upswing which ended in February 1985 has been particularly puzzling. It is often believed to have been the outcome of a "bubble" (G.W. Evans (1986)), possibly brought about by a progressive loss of confidence by portfolio managers in the fundamentals (J.A. Frankel, K.A. Froot (1990.a, 1990.b)). In this paper we provide an economic interpretation of the "long swings" of the dollar identified by C. Engel, J.D. Hamilton (1990) and suggest that economic agents were reacting to fundamentals throughout the 1980s.

We analyse the relevance of the fundamentals with the help of a dynamic model which combines some properties of an error correcting structure with those of the "news" approach. We try to model in this way the behaviour of the dollar/sterling exchange rate over the period 1979-1989. A structural shift has occurred in February-March 1985: over the 1979-1985 subperiod economic agents seem to have reacted to a reduced set of fundamentals only, a finding which might corroborate the "irrational behaviour" interpretation of the dollar upswing set out by P. Krugman (1985) and S. Marris (1985). In the 1985-1989 subperiod economic agents have associated the equilibrium value of the dollar/sterling exchange rate with the behaviour of a larger (more balanced) set of fundamentals.

Section I outlines the main characteristics of the model under investigation, while section II presents the model estimates. Section III analyses the timing and the nature of the structural shifts and section IV concludes the paper.²

² The author would like to thank A. Calamanti, F. Casprini, F. Cesarano M. Grassini and P.L. Rizzi for their useful suggestions.

I. A Simple Model of Exchange Rate Determination

As usual the deviations from purchasing power parity are given by

$$1) d_t = s_t - P_t^b + P_t^u$$

where s_t is the logarithm of the spot exchange rate between the currencies of the countries b and u, P_t^b and P_t^u are the logarithms of the price level in these two countries. d_t can also be defined as the logarithm of the real exchange rate between the currencies concerned.

Money demand is defined as

$$2) P_t^k = a^k m_t^k - b^k y_t^k + c^k i_t^k + G^k, \quad k = b, u$$

where m_t^k , y_t^k and i_t^k are the logarithm of the money supply, the logarithm of real output and the nominal interest rate in the two countries concerned. Domestic and foreign interest rate are linked through the uncovered interest parity condition

$$3) i_t^b - i_t^u = E(s_{t+1}|I_t) - s_t$$

where $E(.)$ denotes the market's expectation conditional upon the market's period t information set I_t . The real interest rate r_t^k is related to the nominal interest rate and to the expected inflation rate through the Fisher interest rate condition

$$4) r_t^k = i_t^k - E(\Delta P_{t+1}^k | I_t), \quad k = b, u.$$

Equations 5), 6) and 7) describe the evolution over time of money supply, real output and real interest rate in the two countries concerned

$$5) m_t^k = e_0^k + e_1^k m_{t-1}^k + e_2^k \Delta m_{t-1}^k + e_3^k \Delta y_{t-1}^k + e_4^k \Delta r_{t-1}^k + v_t^k$$

$$6) y_t^k = d_0^k + d_1^k y_{t-1}^k + d_2^k \Delta m_{t-1}^k + d_3^k \Delta y_{t-1}^k + d_4^k \Delta r_{t-1}^k + w_t^k$$

$$7) r_t^k = h_0^k + h_1^k r_{t-1}^k + h_2^k \Delta m_{t-1}^k + h_3^k \Delta y_{t-1}^k + h_4^k \Delta r_{t-1}^k + l_t^k,$$

where v_t^k , w_t^k and l_t^k are independent, serially uncorrelated shocks with zero mean and constant variance.

Applying a methodology developed by G.C. Chow (1983), we obtain the following dynamic relationship as model solution (the details are set out in appendix I)

$$8) \Delta s_t = \frac{1}{c} \{s_{t-1} - [a m_{t-1}^d - b y_{t-1}^d - r_{t-1}^d + (G^d + E(d_t | I_{t-1}))]\}$$

$$+ \psi_1 [m_t^d - E(m_t^d | I_{t-1})] + \psi_2 [y_t^d - E(y_t^d | I_{t-1})]$$

$$+ \psi_3 [r_t^d - E(r_t^d | I_{t-1})] + e_{8t} + (R_1 - \frac{1+c}{c}) e_{8t-1},$$

where $m_t^d = m_t^b - m_t^u$, $y_t^d = y_t^b - y_t^u$ and $r_t^d = r_t^b - r_t^u$.

The hypothesis that d_t be constant seems to be unrealistic in the time period concerned. We assume instead that the real exchange rate is linked to the cumulated trade balances of the two countries b and u,

$$9) d_t = \tau_0 - \tau_1 (CTB_{t-1}^b - CTB_{t-1}^u) + u_t,$$

where CTB_t^k is the cumulated trade balance in country k and u_t is a residual.¹ Equation 8) can then be rewritten - replacing

r_{t-1}^d by its determinants according to the Fisher relationship - as:

$$10) \Delta s_t = \frac{1}{c} \{s_{t-1} - [am_{t-1}^d - by_{t-1}^d - i_{t-1}^d + E(\Delta P_t^d | I_{t-1}) - \tau_1 CTB_{t-1}^d$$

$$+ \phi] + \psi_1[m_t^d - E(m_t^d | I_{t-1})] + \psi_2[y_t^d - E(y_t^d | I_{t-1})]$$

$$+ \psi_3[r_t^d - E(r_t^d | I_{t-1})] + e_{10t} + (R_1 - \frac{1+c}{c})e_{10t-1},$$

where

$$\phi = G^d + \tau_0, CTB_t^d = CTB_t^b - CTB_t^u, i_t^d = i_t^b - i_t^u \text{ and } P_t^d = P_t^b - P_t^u.$$

The first term on the right hand side is the residual of equation 11) below, which can be interpreted as a cointegration relationship between the exchange rate and money, real output, nominal interest rate, expected inflation rate and cumulated trade balance (differences) in the two countries concerned,

$$11) s_t = am_t^d - by_t^d - i_t^d + E(\Delta P_{t+1}^d | I_t) - \tau_1 CTB_t^d + \phi + e_{11t}.$$

The cointegration regression can be seen as a by-product of the market beliefs on the nature of the factors which determine the (long run) equilibrium value of the exchange rate. A regime shift should bring about a shift in the cointegrating regression and a change in its specification if it alters the market beliefs as to the relative relevance of these factors.

Equation 10) combines some properties of an error correcting model with those of the "news" approach and describes the short run behaviour of the exchange rate. The lagged residual of equation 11) can be seen as the result of the progressive learning by economic agents on the economic policy changes brought about by the U.S. authorities, changes which are

accompanied by shifts in the cointegrating regression.²

II. Estimates of the Exchange Rate Relationships

We have investigated at first the behaviour of the exchange rate between the dollar and the sterling over the period April 1979-June 1989.³ The model examined above suggests that we analyse equation 12) as possible cointegration regression where, in line with J. A. Frankel (1979), we have proxied the expected inflation rate by a long run interest rate:

$$12) s_t = \gamma_0 + \gamma_1 m_t^d + \gamma_2 y_t^d + \gamma_3 i_t^d + \gamma_4 Li_t^d + \gamma_5 CTB_t^d + \gamma_6 t + e_{12t}.$$

Li_t^d is the difference between the long term interest rate in the two countries concerned. Following B.E. Hansen, P.C.B. Phillips (1990), we have added a time trend to the set of regressors, since the m_t^d time series can be modelled as a random walk with drift. As shown in appendix II below, the remaining time series can be modelled as random walks without drift. The CRDW and DF statistics set out in table 1 suggest that the null hypothesis of no cointegration has to be rejected, since the cointegration residual is stationary. (The $\hat{Z}_t$ statistic, however, is not significant at the 5% level and supports the no cointegration hypothesis.) The coefficients of the regression estimates have -with one exception, γ_3 - the signs suggested by the theory.

In order to estimate equation 10) we have to model the "news" about the fundamental variables. The expected values of the impulses have been computed as the fitted values of the OLS estimates of equations 5), 6) and 7), modified in order

to include the lagged residual of the cointegrating regression $\hat{e}_{12t-1}$ as additional regressor.⁴ From an economic point of view s_t , m_t^d , y_t^d and r_t^d may well be determined simultaneously. As a consequence m_t^d , y_t^d and r_t^d and thus the regressors $[m_t^d - \hat{m}_t^{d0}]$, $[y_t^d - \hat{y}_t^{d0}]$ and $[r_t^d - \hat{r}_t^{d0}]$ will be correlated with the residuals of equation 10). As shown in a previous paper (G. Cifarelli (1990)),

TABLE 1

1979:4-1989:6

$$12) s_t = \gamma_0 + \gamma_1 m_t^d + \gamma_2 y_t^d + \gamma_3 i_t^d + \gamma_4 Li_t^d + \gamma_5 CTB_t^d + \gamma_6 t + e_{12t}$$

γ_0	γ_1	γ_2	γ_3	γ_4	γ_5	γ_6	$\bar{R}^2$	S.E.	CRDW	DF	$\hat{Z}_t^*$
7.50	1.53	-1.20	0.16	0.58	-0.0000030	-0.02 ^h	0.87	0.0682	0.51	-4.75	-4.83

1979:5-1989:6

$$13) \Delta s_t = \theta_0 + \theta_1 \hat{e}_{12t-1} + \theta_2 [\hat{m}_t^d - \hat{m}_t^{d0}] + \theta_3 [\hat{y}_t^d - \hat{y}_t^{d0}] + \theta_4 [\hat{r}_t^d - \hat{r}_t^{d0}]$$

$$+ (\theta_{13t} + \theta_2 \hat{z}_{1t} + \theta_3 \hat{z}_{2t} + \theta_4 \hat{z}_{3t}) + \theta_5 e_{13t-1}$$

θ_0	θ_1	θ_2	θ_3	θ_4	θ_5	$\bar{R}^2$	S.E.	D.W.	AR(1)	LM(5) ^{&}
1) -0.002	-0.15	0.95	0.38	-1.50	-0.02	0.09	0.0352	1.99	1.95	5.84
(-0.73) (-3.12) (2.16) (0.52) (-1.70) (-0.26)										
2) -0.002	-0.15	0.95	0.38	-1.38	-	0.09	0.0352	2.04	2.35	6.02
(-0.78) (-2.31) (2.38) (0.59) (-1.64) [∞] -										
3) -	-0.15	0.95	-	-1.46	-	0.11	0.0350	1.99	2.95	6.20
- (-2.33) (2.33) - (-1.77) [∞] -										

* THE NUMBER OF AUTOCORRELATIONS IN THE RESIDUAL SUM VARIANCE ESTIMATOR IS

3. CRITICAL VALUES AT THE 5% LEVEL OF SIGNIFICANCE: $\hat{Z}_t = -5.03$, CRDW = 0.39,

DF = -4.58.⁵

^h WE DO NOT REPORT THE t-RATIOS, WHICH ARE BIASED.

[&] AR(N) = CHI-SQUARE TEST FOR Nth ORDER ARCH, LM(N) = LAGRANGE MULTIPLIER TEST FOR Nth ORDER ERROR STRUCTURE.

[∞] THE t-RATIOS ARE COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

this consistency problem can be solved with the standard instrumental variables approach.⁶

The following relationship has been estimated with OLS:

$$13) \Delta s_t = \theta_0 + \theta_1 \hat{e}_{12t-1} + \theta_2 [\hat{m}_t^d - \hat{m}_t^{d0}] + \theta_3 [\hat{y}_t^d - \hat{y}_t^{d0}]$$

$$+ \theta_4 [\hat{r}_t^d - \hat{r}_t^{d0}] + (e_{13t} + \theta_2 \hat{z}_{1t} + \theta_3 \hat{z}_{2t} + \theta_4 \hat{z}_{3t}) + \theta_5 e_{13t-1}$$

$$\text{where } \hat{e}_{12t-1} = s_{t-1} - (\hat{\gamma}_0 + \hat{\gamma}_1 m_{t-1}^d + \hat{\gamma}_2 y_{t-1}^d + \hat{\gamma}_3 i_{t-1}^d + \hat{\gamma}_4 Li_{t-1}^d$$

$$+ \hat{\gamma}_5 CTB_{t-1}^d + \hat{\gamma}_6 t), \hat{z}_{1t} = m_t^d - \hat{m}_t^d, \hat{z}_{2t} = y_t^d - \hat{y}_t^d \text{ and } \hat{z}_{3t} = r_t^d - \hat{r}_t^d.$$

The empirical findings are reported in table 1. The sign of the significant coefficients in the estimates of equations 13.1), 13.2) and 13.3) is in line with the rationale of the model: an (unexpected) increase in the supply of money brings about a decrease in the value of the domestic currency, whereas an increase in the real interest rate brings about an increase in the value of the domestic currency.⁷ The coefficient of $\hat{e}_{12t-1}$ is significant and negative: the market seems to have a stabilising reaction to a deviation of the exchange rate from its long run value given by the fitted value of the cointegration regression.⁸ The multiple correlation coefficient is rather low. A major weakness of these estimates, however, is that they do not seem to have been stable over the period 1979-1989. A standard Chow test for parameter stability suggests that there has been a shift in the exchange rate relationship in February-March 1985.⁹ As a consequence we have reestimated the relevant relationships over the 1979-1985 and over the 1985-1989 time periods.

The estimates of the period April 1979-March 1985 are reported in table 2 and the estimates of the period February

TABLE 2

1979:4-1985:3

$$12') s_t = \gamma_0 + \gamma_1 m_t^d + \gamma_2 y_t^d + \gamma_3 i_t^d + \gamma_4 Li_t^d + \gamma_5 CTB_t^d + \gamma_6 t + \varepsilon_{12't}$$

	γ_0	γ_1	γ_2	γ_3	γ_4	γ_5	γ_6	$\bar{R}^2$	S.E.	CRDW	DF	$\bar{Z}_t^*$
1)	12.52	2.75	-0.12	-0.014	-	-	-	0.93	0.0600	0.92	-4.55	-4.58
2)	12.33	2.71	-0.10	-0.009	-0.04	-	-	0.93	0.0604	0.90	-4.53	-4.56

1979:5-1985:3

$$13') \Delta s_t = \theta_0 + \theta_1 \hat{\varepsilon}_{12't-1} + \theta_2 [\hat{m}_t^d - \hat{m}_t^{d0}] + \theta_3 [\hat{y}_t^d - \hat{y}_t^{d0}] + \theta_4 [\hat{i}_t^d - \hat{i}_t^{d0}]$$

$$+ (\theta_{13't} + \theta_{22't} + \theta_{32't} + \theta_{42't}) + \theta_5 \varepsilon_{13't-1}$$

	θ_0	θ_1	θ_2	θ_3	θ_4	θ_5	$\bar{R}^2$	S.E.	D.W.	AR(1)	LM(5)
1)	-0.01	-0.23	1.44	0.28	-0.02	-0.08	0.20	0.0334	1.75	3.95	7.39
	(-1.77) (-3.58) (2.68) (0.50) (-1.73) (-0.59)										
2)	-0.01	-0.25	1.56	-	-0.023	-	0.22	0.0329	1.81	2.71	8.86
	(-1.86) (-2.50) (3.43) - (-2.47) ^W -										
3)	-	-0.24	1.56	-	-0.023	-	0.23	0.0334	1.73	3.81	8.82
	- (-2.54) (3.17) - (-2.35) ^W -										

* CRITICAL VALUES AT THE 5% LEVEL OF SIGNIFICANCE: CRDW = 0.78, AND DF = -4.35, -4.76, $\bar{Z}_t = -4.11$, -4.45 (WITH 3 AND 4 REGRESSORS RESPECTIVELY).

^W THE t-RATIOS ARE COMPUTED WITH HETEROSKEDASTICITY CONSISTENT STANDARD ERRORS.

1985-June 1989 are reported in table 3. (The corresponding integration analyses of the relevant time series are set out in appendix II.)

The null of no cointegration is strongly rejected by the data

in both time periods.¹⁰ The overall quality of fit is rather good, especially in the subperiod February 1985-June 1989.

TABLE 3

1985:2-1989:6

$$12'') s_t = \gamma_0'' + \gamma_1'' m_t^d + \gamma_2'' y_t^d + \gamma_3'' i_t^d + \gamma_4'' Li_t^d + \gamma_5'' CTB_t^d + \gamma_6'' t + \varepsilon_{12''t}$$

	γ_0''	γ_1''	γ_2''	γ_3''	γ_4''	γ_5''	γ_6''	$\bar{R}^2$	S.E.	CRDW	DF	$\bar{Z}_t^*$
1)	3.74	0.94	-	0.17	0.32	-0.000003	-0.02	0.84	0.0474	0.86	-5.93	-5.97
2)	6.18	1.40	-	0.32	0.21	-0.000002	-	0.83	0.0499	0.99	-5.74	-5.80

1985:3-1989:6

$$13'') \Delta s_t = \theta_0'' + \theta_1'' \hat{\varepsilon}_{12''t-1} + \theta_2'' [\hat{m}_t^d - \hat{m}_t^{d0}] + \theta_3'' [\hat{y}_t^d - \hat{y}_t^{d0}] + \theta_4'' [\hat{i}_t^d - \hat{i}_t^{d0}]$$

$$+ (\theta_{13''t} + \theta_{22''t} + \theta_{32''t} + \theta_{42''t}) + \theta_5'' \varepsilon_{13''t-1}$$

	θ_0''	θ_1''	θ_2''	θ_3''	θ_4''	θ_5''	$\bar{R}^2$	S.E.	D.W.	AR(1)	LM(5)
1)	0.01	-0.42	0.71	-1.89	0.02	-0.09	0.37	0.0334	1.63	0.85	7.19
	(1.94) (-4.37) (1.87) (-2.43) (1.52) (-0.64)										
2)	0.01	-0.45	0.72	-1.63	0.03	-	0.37	0.0335	1.78	1.15	7.49
	(1.57) (-4.72) (1.98) (-2.86) (2.25) -										

* CRITICAL VALUES AT THE 5% LEVEL OF SIGNIFICANCE: CRDW = 0.78, DF = -4.76, $\bar{Z}_t = -4.74$ (EQUATION 12'' 1), -4.45 (EQUATION 12'' 2)).

The significant coefficients of equations 13'1), 13'2), 13'' 1) and 13'' 2), and all the coefficients of equations 13'3) have the appropriate signs. In the first subperiod the exchange rate seems to have been unaffected by "news" on the relative real output behaviour, whereas, after the February 1985 shift, it is the "news" about the relative real interest rate behaviour that have become irrelevant as determinants of exchange rate behaviour. In both subperiods the coefficients of the lagged equilibrium error suggest a strong equilibrating reaction of the

market to a deviation of the exchange rate from its long run value.

It should be noticed, however, that there has been a change in the specification of the cointegrating regression. Over the period February 1985-June 1989 the exchange rate appears to be cointegrated with the money supply, the nominal interest rate, the expected inflation rate and the cumulative trade balance (differences) in the two countries concerned. (Cointegration holds also if we do not include a time trend among the regressors.) All the coefficients have - with one exception, γ_3'' - the appropriate signs. Over the period April 1979-March 1985 portfolio managers have associated the long run (equilibrium) value of the exchange rate with the behaviour of a reduced set of fundamentals only: money supply, real output and nominal interest rate in the two countries concerned. (The estimates of equation 12'2) suggest that the expected inflation differential is irrelevant for cointegration.) This finding provides a possible explanation for the sharp devaluation of the dollar of February 1985. The perception of a shift in the reaction of the federal authorities to a growing deficit of the U.S. balance of payments (away from a policy of "benign neglect") has brought about a rapid decline of what the market believed to be the equilibrium value of the exchange rate. The existing cointegrating regression - which did not include the cumulative trade balances as regressors - provided an equilibrium value of the exchange rate which was no longer correct.¹¹ It seems that the dollar shift of February 1985 was not caused by a shift in the fundamentals, but by a change in what the market believed to be the set of factors which determined the exchange rate

behaviour. This hypothesis is further corroborated by the analysis of the timing of the shifts of the exchange rate change relationship and of the fundamental variables relationships.

III. The Tests for Structural Change

We have applied the version of the Quandt log-likelihood approach (R.E. Quandt (1958, 1960)) set forth by R.L. Brown, J. Durbin and J.M. Evans (1975) in order to assess directly from the data if and when a structural change has occurred in the relationships concerned.¹² Quandt's approach involves the computation of the following log-likelihood ratio test

$$\lambda_r = r \log \hat{\sigma}_0 + (T-r) \log \hat{\sigma}_1 - T \log \hat{\sigma}$$

where $\hat{\sigma}_0$ and $\hat{\sigma}_1$ are estimated standard errors of the old and new regimes, $\hat{\sigma}$ is the standard error over the whole sample, T is the total number of observations and r is the last date of the old regime. The estimate of the point at which the switch from one specification to another has occurred is assessed by computing λ_r for all possible r and then choosing the value of r at which λ_r attains its minimum.¹³

The λ_r statistics plotted in table 4 suggest that equation 13.2) shifts in February-March 1985. This equation seems to have been stable in the 1985-1989 subperiod and rather unstable in the 1979-1985 subperiod. The λ_r statistics plotted in table 5 suggest that equation 7) has been unstable from the beginning of 1981 to the end of 1982, the most likely date for a shift being March-April 1982. This shift coincides with the

TABLE 4

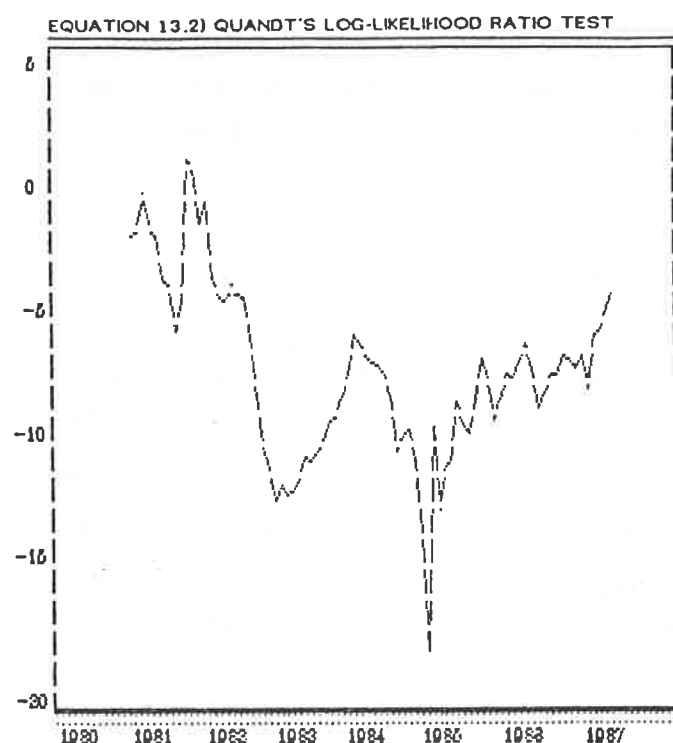
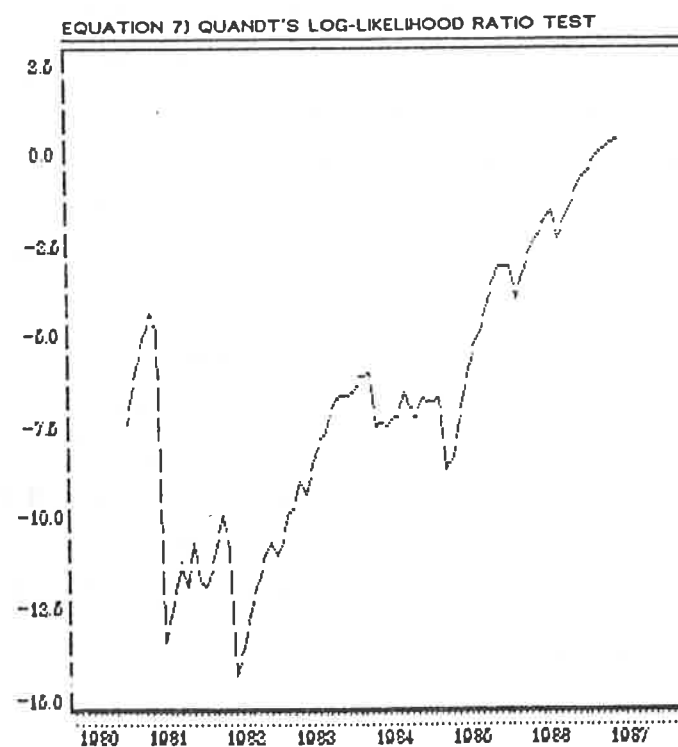


TABLE 5



peak of the short term U.S.-U.K. real interest rate differential and precedes an U.S. real interest rate structural shift. (The latter is believed by T. Kinal, K. Lahiri (1988) to have occurred at the beginning of 1983, following a change in the Federal Reserve operating procedures.) Surprisingly, the relative money supply and relative real output relationships have been stable in the time period concerned: the likelihood surfaces of equations 5) and 6) are smooth, with no discernible minima, and are not reported here.

The λ_r graphs corroborate the hypothesis that the sharp decline of the exchange rate of February 1985 was not caused by a shift in the fundamentals. The exchange rate relationship shifts three years after the relative real interest rate relationship, a result which is compatible with a Bayesian learning interpretation a' la Lewis (K.K. Lewis (1989.a, 1989.b)). The persistence in 1983 and 1984 of the dollar appreciation might well be due to the slow and gradual perception by portfolio managers of this real interest rate regime shift and/or of its causes.¹⁴ Indeed, the cointegration analysis of the previous section suggests that they associated the long run exchange rate level with nominal rather than with real interest rate behaviour. The February-March 1985 shift was probably brought about by the change in the specification of the cointegrating regression examined in the previous section.

IV. Conclusion

In this paper we have provided an interpretation in terms of fundamental variables of the dollar behaviour in the 1980s. Economic agents seem to have reacted throughout to

fundamental variables, adjusting the exchange rate to a long run equilibrium value provided by a cointegration relationship which includes the fundamental variables. The dollar upswing and the subsequent sharp decline of 1985 have been caused by the incorrect specification of the cointegrating regression, i.e. by an incorrect understanding of the mechanisms which determined the equilibrium value of the dollar. Portfolio managers were led to associate the equilibrium value of the exchange rate with the behaviour of an incomplete set of fundamentals, neglecting those fundamentals which tended to reduce the expected (equilibrium) value of the dollar. The sudden realisation of this error brought about a change in the specification of the cointegrating relationship and a fall in what the market believed to be the equilibrium value of this currency.

Appendix I. Model Solution

Consider the following relationship, which describes the exchange rate behaviour:

$$A.1) s_t = \alpha Z_t + \beta E(s_{t+1}|I_t) + e_t$$

where Z_t is a vector of exogenous variables and e_t is a normal, serially uncorrelated residual. Following the methodology set out by G.C. Chow (1983), we assume that equation A.1) is consistent with the relationship

$$A.2) s_t = R_0 e_t + R_1 e_{t-1} + \dots + R_n e_{t-n} + k_0 Z_t + k_1 Z_{t-1} + \dots + k_n Z_{t-n}.$$

Advancing the time subscripts in A.2) by 1, the corresponding one period ahead forecast errors are given by:

$$A.3) s_{t+1} - E(s_{t+1}|I_t) = R_0 e_{t+1} + R_1 e_t + k_0 [Z_{t+1} - E(Z_{t+1}|I_t)].$$

This equation can be used to replace $E(s_{t+1}|I_t)$ in equation A.1) by s_{t+1} minus a linear combination of e_{t+1} , e_t and $[Z_{t+1} - E(Z_{t+1}|I_t)]$. Assuming that $R_0 = 1$, we obtain:

$$A.4) \Delta s_{t+1} = \left(\frac{1-\beta}{\beta}\right) [s_t - \frac{\alpha}{1-\beta} Z_t] + (R_1 - \frac{1}{\beta}) e_t + e_{t+1} + k_0 [Z_{t+1} -$$

$$E(Z_{t+1}|I_t)],$$

where $s_t - \frac{\alpha}{1-\beta} Z_t$ can be seen as the lagged residual of a (first stage) cointegrating regression.

If we want to apply this methodology to the model above, we have to solve it in terms of a relationship analogous to equation A.1). Taking conditional expectations of equation 1) we obtain, rearranging terms:

$$E(s_{t+1}|I_t) = E[(P_{t+1}^b - P_{t+1}^u)|I_t] + E(d_{t+1}|I_t)$$

and, using equation 3), we have:

$$A.5) s_t = E[(P_{t+1}^b - P_{t+1}^u)|I_t] + E(d_{t+1}|I_t) - (i_t^b - i_t^u).$$

Using the Fisher relationships, we obtain:

$$A.6) s_t = E(d_{t+1}|I_t) - (r_t^b - r_t^u) + (P_t^b - P_t^u).$$

With the help of equations 2) and 4) we get:

$$A.7) P_t^k = \left[\frac{1}{1+c^k} \right] [a^k m_t^k - b^k y_t^k + c^k r_t^k + G^k] + \left[\frac{c^k}{1+c^k} \right] E(P_{t+1}^k | I_t),$$

$k = b, u.$

Substituting back into equation A.6), assuming that $P_t^d = P_t^b - P_t^u$, $m_t^d = m_t^b - m_t^u$, $y_t^d = y_t^b - y_t^u$, $r_t^d = r_t^b - r_t^u$ and $G^d = G^b - G^u$ and that $a = a^b = a^u$, $b = b^b = b^u$ and $c = c^b = c^u$, we obtain:

$$A.8) s_t = \left[\frac{1}{1+c} \right] [a m_t^d - b y_t^d - r_t^d + G^d] + \left[\frac{c}{1+c} \right] E(P_{t+1}^d | I_t)$$

$$+ E(d_{t+1} | I_t).$$

Given that $E(P_{t+1}^d | I_t) = E(s_{t+1} | I_t) - E(d_{t+1} | I_t)$, we can rewrite equation A.8) as:

$$A.9) s_t = \left[\frac{1}{1+c} \right] [a m_t^d - b y_t^d - r_t^d + G^d] + \left[\frac{c}{1+c} \right] E(s_{t+1} | I_t),$$

$$\text{where } G^{d'} = G^d + E(d_{t+1} | I_t).$$

Equation A.9), with the addition of a stochastic residual, e_t , is analogous to equation A.1). With the help of the Chow methodology it can be solved as equation 8) of the text. (The latter is reported, however, lagged by one time period.)

Appendix II. Integration Analysis

As usual in cointegration analysis, we pretest the relevant time series for unit roots. We apply the univariate unit root test statistics set forth by P.C.B. Phillips, P. Perron (1988). As we see in the tables below, the time series s_t , y_t^d , i_t^d , Li_t^d and CTB_t^d appear to be $I(1)$ over the three time periods of interest (they can be modelled as random walks with no drift). The m_t^d

time series can be modelled as a random walk with drift over the periods February 1985-June 1989 and April 1979-June 1989

TABLE A.1

1979:4-1989:6

	s_t	m_t^d	y_t^d	i_t^d	Li_t^d	CTB_t^d	5 % CRIT. VAL.*
$Z(t_{\alpha})$	-1.25	2.41	-1.84	-2.56	-2.37	-0.44	-2.89
$Z(\#_1)$	1.04	12.59	2.08	3.28	2.85	1.20	4.71
$Z(t_{\alpha})$	-1.16	-1.43	-2.21	-2.90	-2.16	-2.36	-3.45
$Z(\#_2)$	0.70	9.89	1.89	2.81	2.02	2.75	4.88
$Z(\#_3)$	0.80	5.14	2.45	4.21	2.99	2.98	6.49

* UNDER THEIR RESPECTIVE NULL HYPOTHESES $Z(\#_3)$, $Z(\#_2)$, $Z(\#_1)$, $Z(t_{\alpha})$ AND $Z(t_{\alpha})$ WILL HAVE THE $\#_3$, $\#_2$, $\#_1$, τ_{μ} AND τ_{τ} DISTRIBUTIONS (W.A. FULLER (1976), D.A. DICKEY, W.A. FULLER (1981)).

TABLE A.2

1979:4-1985:3

	s_t	m_t^d	y_t^d	i_t^d	Li_t^d	CTB_t^d	5% CRIT. VAL.
$Z(t_{\alpha})$	0.01	-0.49	-1.59	-2.45	-1.60	2.64	-2.93
$Z(\#_1)$	1.21	2.30	1.45	3.03	2.00	4.31	4.86
$Z(t_{\alpha})$	-2.97	-2.90	-1.64	-2.42	-3.43	1.20	-3.50
$Z(\#_2)$	4.31	4.73	1.05	2.11	4.31	2.84	5.13
$Z(\#_3)$	5.10	5.11	1.38	3.13	5.93	3.46	6.73

TABLE A.3

1985:2-1989:6

	s_t	m_t^d	y_t^d	i_t^d	Li_t^d	CTB_t^d	5% CRIT. VAL.
$Z(t_{\alpha})$	-2.64	0.24	-1.32	-2.32	-1.97	-2.33	-2.93
$Z(\#_1)$	4.22	10.97	0.95	2.75	2.12	3.80	4.86
$Z(t_{\alpha})$	-1.41	-3.41	-2.64	-2.23	-1.93	-0.47	-3.50
$Z(\#_2)$	2.78	10.17	3.33	2.02	1.41	2.52	5.13
$Z(\#_3)$	3.45	5.98	4.90	2.96	1.92	2.71	6.73

and as a random walk without drift over the period April 1979 - March 1985.

A preliminary investigation showed that the s_t time series might well be affected - in February 1985 - by both a change in trend and a shift in its level. In this case the standard tests of table A.1 will not be able to reject the unit root hypothesis even if the true data generating mechanism is that of stationary fluctuations around a trend. We have therefore applied the "Model C" version of the two step Perron unit root test (P. Perron (1989)) which allows - under both the null and alternative hypotheses - for the presence of a one time change in the level and in the slope of the trend function.

This test is performed by applying the standard Dickey-Fuller procedure to the residuals of the following first stage regression

$$s_t = \mu + \beta t + \gamma DU_t + \theta DT_t + e_t,$$

where $DT_t = t$ if $t > TB$, 0 otherwise, $DU_t = 1$ if $t = TB + 1$, 0 otherwise and TB refers to the time of break (February 1985). We obtain the following OLS estimates of $\tilde{e}_t = \alpha \tilde{e}_{t-1} + \psi \Delta \tilde{e}_{t-1} + u_t$:

α	ψ	$\bar{R}^2$	S.E.	D.W.	DF	5% CRIT. VAL.
0.87	0.03	0.72	0.0321	2.00	-2.48	-4.24
(16.51)	(0.30)					

The null hypothesis of a unit root cannot be rejected over the period 1979-1989, since the DF statistic is not significant at the 5% level, with a ratio of pre-break sample size to total sample size of 0.56. This finding is in line with the results of the unit roots tests of table A.1 above.

Notes

(1) We are following here P. Hooper, J. Morton (1982) and V.S. Somanath (1986). If TB_t^k is the trade balance of country k at time t , $CTB_t^k = \sum_{i=0}^t TB_{t-i}^k$.

(2) The absorption of great innovations, which cannot be evaluated with the help of past experience, is slow and can be modelled better if we abandon rational expectations in favour of an alternative hypothesis on expectations formation. An analysis of this problem is set out in G. Cifarelli (1990).

(3) Data are sampled monthly. The spot dollar sterling rate is reported in the "Financial Times". Money supply (seasonally adjusted), industrial output (seasonally adjusted), the trade balance and consumer price time series in the U.S. and in the U.K. are published in the O.E.C.D. "Main Economic Indicators". The British interbank interest rate and the U.S. federal funds interest rate time series are reported in the O.E.C.D. "Financial Statistics". The yield of the long-term U.S. Government bonds and of the U.K. $2\frac{1}{2}$ per cent consols time series (the expected inflation rate proxies) are published in the O.E.C.D. "Main Economic Indicators". The real interest rates are ex post rates. All interest rates are per cent per month.

(4) We have used parsimonious estimates of these relationships, and have dropped the regressors with coefficients not significantly different from zero.

(5) The critical values of the $\hat{Z}_t$ statistic have been tabulated by P.C.B. Phillips, S. Ouliaris (1990). The critical values of the DF and CRDW statistics have been tabulated by R.F. Engle, B.S. Yoo (1987).

(6) We have replaced m_t^d , y_t^d and r_t^d with their respective fitted

values, $\hat{m}_t^d$, $\hat{y}_t^d$ and $\hat{r}_t^d$, obtained by means of OLS regressions of m_t^d , y_t^d and r_t^d on a constant, the lagged values of m_t^d , y_t^d and r_t^d (with lags of up to four months), on $\hat{e}_{12t-1}$, on s_{t-1} , on the current and lagged values (with lags of up to two months) of the difference between the logarithms of the consumer price level and on the difference between the current values of the cumulated trade balances in the two countries concerned.

(7) If the fitted value of equation 7) $\hat{r}_t^{d0}$ is seen - following J. Huizinga, F.S. Mishkin (1984) - as the ex ante real interest rate, $-(\hat{r}_t^d - \hat{r}_t^{d0})$ is the inflation forecast error. An unexpected increase in the inflation rate brings about an unexpected decrease in the real interest rate and a decline in the value of the domestic currency (and vice versa).

(8) The spot exchange rate at time t can be seen - discarding the "news" - as a weighted average of the spot exchange rate at time $t-1$ and of the equilibrium (expected) exchange rate at time $t-1$, provided by the cointegration regression,

$$s_t = (1 - \theta_1)s_{t-1} + \theta_1\hat{s}_{t-1}.$$

Indeed, readjusting terms, this relationship can be rewritten as:

$$\Delta s_t = -\theta_1(s_{t-1} - \hat{s}_{t-1}) = -\theta_1\hat{e}_{t-1}.$$

If, as is the case in the estimates above, θ_1 is positive (and $-\theta_1$ is negative), the spot exchange rate will converge to its equilibrium value, given by the cointegration regression. (A detailed analysis of this relationship can be found in J.A. Frankel, K.A. Froot (1987).)

(9) The stability of equation 13) in the period March 1979 - June 1989 is assessed with the help of the dummy variable D_1 , which postulates a shift in February - March 1985:

$D_1 = 0$ in the period April 1979 - February 1985;

$D_1 = 1$ in the period March 1985 - June 1989.

We estimate the relationship

$$\begin{aligned} \Delta s_t = & \theta_0 + \theta_1\hat{e}_{12t-1} + \theta_2[\hat{m}_t^d - \hat{m}_t^{d0}] + \theta_3[\hat{y}_t^d - \hat{y}_t^{d0}] \\ & + \theta_4[\hat{r}_t^d - \hat{r}_t^{d0}] + \theta_5e_{13t-1} + \theta_6D_1 + \theta_7D_1\hat{e}_{12t-1} + \theta_8D_1[\hat{m}_t^d - \hat{m}_t^{d0}] \\ & + \theta_9D_1[\hat{y}_t^d - \hat{y}_t^{d0}] + \theta_{10}D_1[\hat{r}_t^d - \hat{r}_t^{d0}] + (e_{13t} + \theta_2\hat{z}_{1t} + \theta_3\hat{z}_{2t} + \theta_4\hat{z}_{3t}) \end{aligned}$$

and test the hypothesis $H_0: \theta_6 = \theta_7 = \theta_8 = \theta_9 = \theta_{10} = 0$.

We obtain the following F test estimates: $F_{(5,112)} = 4.46$, significant at the 5% level.

(10) With the help of the Johansen multivariate test (S. Johansen (1988)) we cannot reject the hypothesis that the cointegration regressions 12'1) and 12''1) are unique over the relevant time periods.

(11) From this point of view this paper supports the hypothesis set out by P. Krugman (1985) and S. Marris (1985) that the market had failed to correctly take into account the implications of the increasing U.S. indebtedness for the future value of the dollar. In a previous paper (G. Cifarelli (1992)) we have suggested that there is strong evidence that the dollar appreciation up to 1985 was not caused by a bubble. For an analogous opinion, see P. Minford (1989).

(12) The parsimonious specifications of equations 5), 6) and 7) used in the estimation of the model over the period April 1979 - June 1989 are the following:

$$\begin{aligned} m_t^d &= e_0 + e_1m_{t-1}^d + e_2\Delta m_{t-1}^d + e_3\Delta y_{t-1}^d + e_5\hat{e}_{12t-1} + v_t^d \\ y_t^d &= d_0 + d_1y_{t-1}^d + d_3\Delta y_{t-1}^d + d_4\Delta r_{t-1}^d + d_5\hat{e}_{12t-1} + w_t^d \\ r_t^d &= h_0 + h_1r_{t-1}^d + h_5\hat{e}_{12t-1} + l_t^d. \end{aligned}$$

$$(13)\lambda_r = \log_{10} \left(\frac{\text{maximum likelihood of the observations given } H_0}{\text{maximum likelihood of the observations given } H_1} \right)$$

where H_1 is the hypothesis that the observations in the time

segments $(1, \dots, r)$ and $(r+1, \dots, T)$ come from two different regressions. If H_1 prevails and structural change occurs at some time $t = r^*$, λ_r will attain a minimum value at r^* .

Quandt's approach is conditional on the assumption that the change has occurred all at once. However, the behaviour of the graph of λ_r over time sheds light on the stability of the relationship of interest: a jagged graph suggests that this relationship is unstable, a smooth graph that it is stable.

(14) As pointed out by V. Koromzay, J. Llewellyn and S. Potter (1987), the market had some difficulty in grasping the consequences for the real interest rate (and thus for the exchange rate) of the growing deficit of the U.S. balance of payments and of the anti inflationary stance of the Federal Reserve. This finding is not surprising: the U.S. economy had been subjected in the early 1980s to a number of monetary and fiscal shocks with unexpected (overshooting) effects on inflation, output and budget deficit. (A detailed analysis of these policies can be found in P.C. Roberts (1989).)

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