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IMF WORKING PAPER

WP/90/118

INTERNATIONAL MONETARY FUND

Research Department

Long-Run Purchasing Power Parity and the Dollar-Sterling Exchange Rate in the 1920s

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December 1990

Abstract

This paper replies to Ahking's (1990) re-examination of Taylor and McMahon's (1988) analysis of long-run purchasing power parity in the 1920s. We demonstrate that Ahking's conclusions are only partially correct and re-establish our conclusion that a form of long-run purchasing-power parity did in fact hold for dollar-sterling during this period. The new results are also employed to gauge the degree of overvaluation of sterling relative to the imposed prewar parity of $4.86 upon its return to gold and for 12 months afterwards.

JEL Classification Numbers:
044, 431, 432

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Contents

Summary

A recent paper re-examines an existing analysis of long-run purchasing power parity (PPP) during the 1920s float, concentrating on the dollar-sterling exchange rate. That paper claims to find evidence of deterministic as well as stochastic trends in the time series and argues that, once such trends are removed from the data, the evidence in support of long-run PPP is weakened. This present paper demonstrates that these conclusions are only partially correct.

Britain's decision to return to the gold standard at the prewar parity of $4.86 to the pound was seen by Keynes and others as an overvaluation of sterling by at least 10 percent--which seems to be the received wisdom on the issue. Other commentaries suggest, however, that sterling was either undervalued against the dollar at $4.86 or at least not overvalued quite to the degree suggested by Keynes. The results of econometric analysis in this paper shed some light on this issue.

The calculations in this analysis suggest an overvaluation of sterling against the U.S. dollar on sterling's return to the gold standard of the order of only some 5 percent. Moreover, further analysis suggests that movements in relative prices had moved the equilibrium rate close to the imposed parity by the end of 1925 and that, by mid-1926, sterling was, in fact, undervalued against the dollar by some 2 percent. Although these figures may seem to fly in the face of the received view, they are not in fact widely different from those reported in other recent studies. Some form of long-run purchasing power parity appears to have held between the United States and the United Kingdom for virtually the whole of the 1920s float.

I. Introduction

In a recent paper, Ahking (1990) re-examines Taylor and McMahon's (1988) analysis of long-run purchasing power parity (PPP) during the 1920s float, 1921-1924, concentrating on the dollar-sterling exchange rate. In particular, Ahking claims to find evidence of deterministic as well as stochastic trends in the time series and argues that, once such trends are removed from the data, the evidence in support of long-run PPP is weakened. In this paper we demonstrate, using the latest econometric and statistical techniques, that Ahking's conclusions are only partially correct. In particular, whilst we refute his claim of the presence of a deterministic trend in the dollar-sterling exchange rate, we substantiate the presence of a deterministic trend in U.K. wholesale prices. By incorporating this fact into the econometric analysis, however, we are able considerably to improve upon the original analysis of Taylor and McMahon by demonstrating that some form of long-run purchasing-power parity did in fact hold for the whole of the period from early 1921 until the return to gold in mid-1925. We then go on to use our new results to gauge the degree of overvaluation of sterling relative to the imposed prewar parity of $4.86 both immediately upon its return and for 12 months afterwards.

The remainder of the paper is set out as follows. Rather than proceeding directly to the econometric analysis, in Section II we derive some initial conjectures concerning the exchange rate and prices simply from an examination of a graph of the relevant time series. Section III contains the main statistical and econometric analysis and results while Section IV uses our new results to examine the question of the over-valuation of sterling on its return to the Gold Standard. A final section concludes.

II. An Informal Look at the Data

Before embarking on the formal statistical and econometric analysis, it is worthwhile visually examining the general pattern of the series under consideration, ie the dollar-sterling exchange rate and U.K. and U.S. wholesale prices for the period early 1921 until Britain's return to the Gold Standard in mid-1925. 1/ Chart 1 graphs the three time series. At a very informal level, inspection of Chart 1 does not suggest a strong tendency for the three-time series to diverge from one another over the period. Indeed, although their behavior clearly does not accord with continuous purchasing power parity, there does seem to be a tendency for the exchange rate to approximate such a relationship: the steep decline in U.K. prices relative to U.S. prices in 1921 coincides with an appreciation of

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1/ Data sources are the same as in Taylor and McMahon, 1988: the dollar-sterling exchange rate data are from Einzig, 1937 the wholesale price data are from Tinbergen, 1934.

sterling against the dollar; the broad stability of both price series over the next two and a half years is matched by a similar degree of stability in the dollar-sterling exchange rate; and the relative decline in U.K. wholesale prices from the end of 1924 coincides more or less with a sterling appreciation.

Thus, at this purely informal level, we might begin to entertain the following conjectures. First, some form of long-run purchasing power parity held for the dollar-sterling exchange rate over this period. Secondly, Ahking is not justified in excluding data points at the beginning of the sample period in order to obtain his desired results. Thirdly, by the same token, Taylor and McMahon may have been mistaken in excluding data points at the end of the sample period in an attempt to establish long-run purchasing power parity. We shall demonstrate formally that each of these conjectures is justified.

III. Unit Roots, Deterministic Trends and Cointegration

Ahking states (page 914) that necessary but not sufficient conditions for a set of time series to be cointegrated are that each of the time series be integrated of the same order and that the series contain no deterministic components. Neither of these statements is quite correct. On the second issue, as noted informally by Granger (1986) and Engle and Granger (1987, page 259) and more formally by Engle and Yoo (1987), it is quite possible that two or more non-stationary time series may have trends in mean which are similar enough to cancel out one another in the cointegrating relationship. 1/ Further, even if this is not the case, it may well be that a set of time series cointegrate <u>about</u> a linear trend--that is to say, there may be some linear combination which eliminates the stochastic trend but leaves the deterministic trend intact. Since deterministic trends are less worrying among economic relationships than stochastic trends (they are perfectly predictable and can often be explained in economic terms), such a situation may still be of considerable economic interest (see, for example, Johansen and Juselius, 1990). Ahking's results do, however, suggest that a more careful treatment of this issue is warranted than was originally given by Taylor and McMahon.

Perron (1988) demonstrates that if a series is stationary about a linear trend but no allowance for this is made in the construction of the unit root test, then the probability of a type II error (failure to reject

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1/ That is to say, the same linear combination which eliminates the unit root may also eliminate the trend--see Engle and Yoo, 1987.

the unit root hypothesis when it is false) may be high. 1/ Perron suggests the following strategy for testing for unit root behavior in a series y_t. The following regression is estimated by OLS:

$$y_t = \kappa + \lambda(t-T/2) + \delta y_{t-1} + u_t \quad (1)$$

where the sample size is T+1 and u_t may be serially correlated and heterogeneously distributed. 2/ The semi non-parametric test statistics developed by Phillips (1987a,b) and Phillips and Perron (1986) can then be used to test the following hypotheses:

$$H_A:\delta = 1; \quad H_B:(\kappa,\lambda,\delta) = (0,0,1); \quad H_C:(\lambda,\delta) = (0,1)$$

The appropriate test statistics are, in fact, transforms of the standard t-statistic for H_A and of the standard F-statistics for H_B and H_C (and we denote them $Z(\tau_\tau)$, $Z(\Phi_2)$ and $Z(\Phi_3)$, respectively). If the unit root hypothesis can be rejected at this juncture, there is no need to proceed. If it cannot, however, then greater test power may be obtained by estimating the regression

$$y_t = \tilde{\kappa} + \tilde{\delta} y_{t-1} + \tilde{u}_t \quad (2)$$

and testing the hypotheses

$$H_D:\tilde{\delta}=1 \quad \text{and} \quad H_E:(\tilde{\kappa},\tilde{\delta}) = (0,1)$$

using the Phillips-Perron transforms of the relevant t-statistic and F-statistic ($Z(\tau_\tau)$ and $Z(\Phi_1)$). This is only valid, however, if the drift term in (1), κ, is zero since $Z(\tau_\mu)$ and $Z(\Phi_1)$ are not invariant with respect to κ. Thus, the statistics $Z(\tau_\mu)$ and $Z(\Phi_1)$ should only be used to provide additional evidence on the unit root hypothesis if the value of $Z(\Phi_2)$ suggests that H_B cannot be rejected (see Perron, 1988).

Table 1 contains results of the full set of Phillips-Perron statistics suggested in the preceding discussion, for each of the time series under investigation. Unlike Ahking, we do not arbitrarily exclude data points at the beginning of the sample but, as a point of comparison with Taylor and

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1/ Alternatively expressed, the test will lack power. The intuition behind Perron's formal proof can be seen as follows. Suppose the true data-generating process is $y_t = \alpha + \beta t + u_t$, where u_t is stationary white noise--i.e., y is stationary about a linear trend. If we estimate the AR(1) model $y_t = \gamma + \rho y_{t-1} + \epsilon_t$ then ρ will be forced to unity, so that the AR(1) model is equivalent to $y_t = y_0 + \gamma t + \tilde{\epsilon}_t$, where $\tilde{\epsilon}_t = \Sigma_0^t \epsilon_t$, which approximates a linear trend.

2/ See Perron (1988) for the precise set of assumptions concerning the error term. The assumptions are sufficiently weak to allow y_t to follow a general ARMA or (subject to the stationarity of the exogenous variables) ARMAX process.

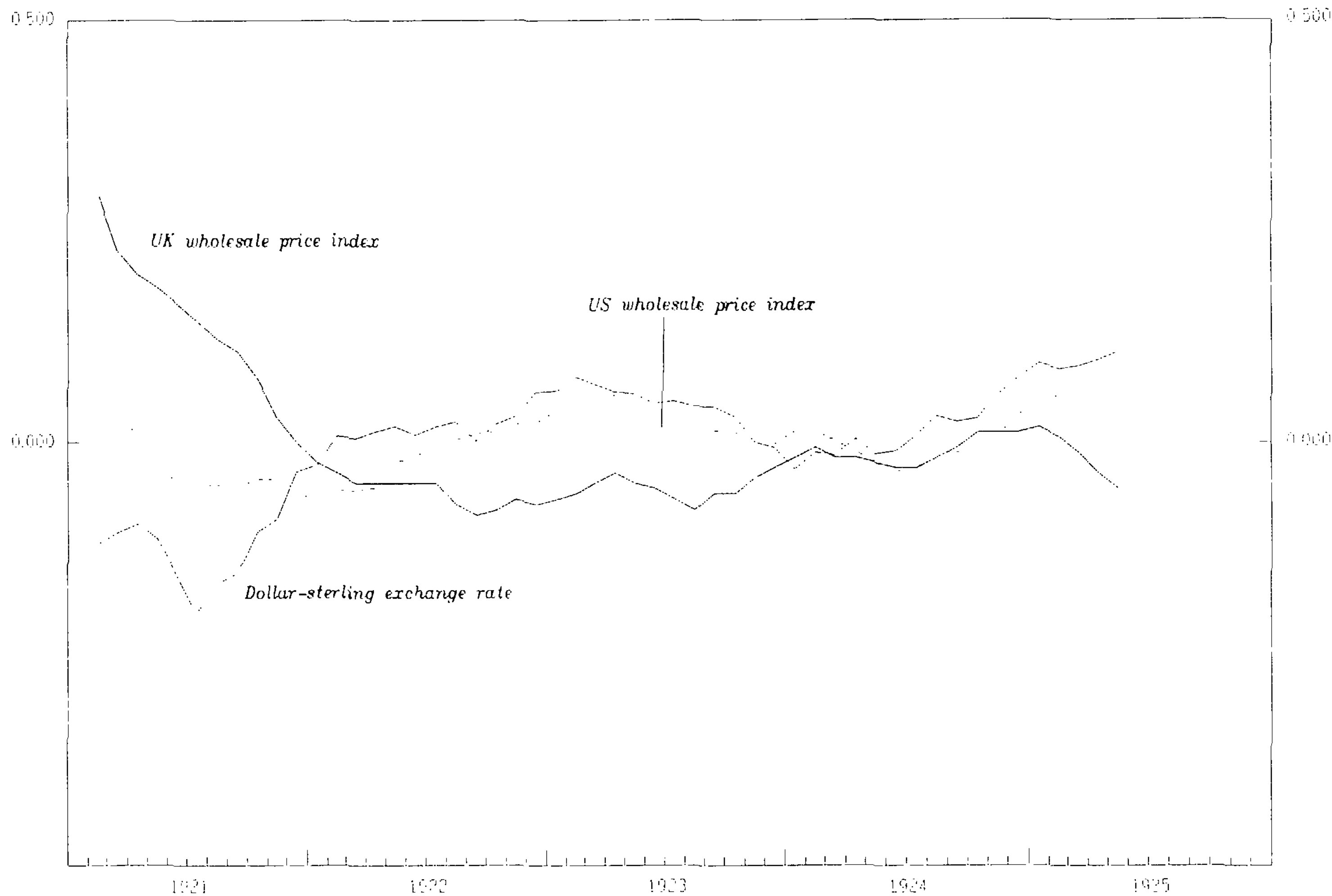
Table 1. Unit Root Tests *

A) Sample period including the last 12 months of the float

Series	Sample period	Lag truncation parameter	Test statistic	Second difference	First difference	Level
Dollar-Sterling	21:1-25:5	2	$Z(\tau_\mu)$:	-13.282	-5.799	-1.245
			$Z(\Phi_1)$:	88.191	16.867	1.647
			$Z(\tau_\tau)$:	-13.144	-5.749	-1.752
			$Z(\Phi_2)$:	57.586	11.052	1.600
			$Z(\Phi_3)$:	86.377	16.577	1.553
UK WPI	21:2-25:5	6	$Z(\tau_\mu)$:	-10.377	-4.344	-4.731
			$Z(\Phi_1)$:	53.301	9.607	14.722
			$Z(\tau_\tau)$:	-10.818	-4.217	-3.979
			$Z(\Phi_2)$:	39.077	6.729	13.235
			$Z(\Phi_3)$:	58.473	9.950	15.013
US WPI	21:2-25:5	1	$Z(\tau_\mu)$:	-10.415	-4.371	-2.093
			$Z(\Phi_1)$:	54.241	9.544	2.246
			$Z(\tau_\tau)$:	-10.476	-4.362	-3.013
			$Z(\Phi_2)$:	36.596	6.429	3.818
			$Z(\Phi_3)$:	54.849	9.615	5.689

CHART 1

US-UK EXCHANGE RATE AND WHOLESALE PRICES, 1921-25 1/



1/ In natural logarithms, mean removed.

Table 1 (Concluded). Unit Root Tests *

B) Sample period excluding the last 12 months of the float

Series	Sample period	Lag truncation parameter	Test statistic	Second difference	First difference	Level
Dollar-Sterling	21:1-24:5	2	$Z(\tau_\mu)$:	-11.661	-5.029	-1.532
			$Z(\Phi_1)$:	68.185	12.744	1.397
			$Z(\tau_\tau)$:	-11.515	-5.160	-1.122
			$Z(\Phi_2)$:	44.311	8.944	0.915
			$Z(\Phi_3)$:	66.415	13.407	1.159
UK WPI	21:2-24:5	6	$Z(\tau_\mu)$:	-10.062	-4.046	-4.948
			$Z(\Phi_1)$:	49.869	8.484	16.781
			$Z(\tau_\tau)$:	-9.929	-4.560	-3.221
			$Z(\Phi_2)$:	32.743	7.832	24.059
			$Z(\Phi_3)$:	48.760	11.380	26.354
US WPI	21:2-24:5	1	$Z(\tau_\mu)$:	-9.464	-3.823	-2.091
			$Z(\Phi_1)$:	44.750	7.258	2.536
			$Z(\tau_\tau)$:	-9.642	-3.685	-2.679
			$Z(\Phi_2)$:	30.977	4.690	2.887
			$Z(\Phi_3)$:	46.461	7.028	3.965

* The null hypotheses and test statistics are discussed in Section II and defined in Perron, 1988. The critical values are as follows (Fuller, 1976, Dickey and Fuller, 1981):

Critical values:	10 percent	5 percent	2.5 percent	1 percent
$Z(\tau_\mu)$	-2.63	-3.00	-3.33	-3.75
$Z(\Phi_1)$	3.78	4.59	5.38	6.43
$Z(\tau_\tau)$	-3.24	-3.60	-3.95	-4.38
$Z(\Phi_2)$	4.03	4.68	5.31	6.09
$Z(\Phi_3)$	5.34	6.25	7.16	8.27

McMahon, we report results both including and excluding the last 12 months of the float--the results are in any case qualitatively identical. In each case we allowed for autocorrelation of the same order as suggested by Ahking. Following the suggestion of Dickey and Pantula (1987) we also tested sequentially for higher-order unit roots by applying the tests to the data in first and second differences in addition to log-levels. The results for the differenced data suggest the absence of higher-order unit roots. At the 5 percent significance level, the only statistic values which are significant for the data in levels are those for the U.K. wholesale price index, irrespective of whether the last 12 months of data are included in the sample. These results thus confirm Ahking's finding for U.K. prices but do not support his argument that the exchange rate is stationary about a trend in mean. As Ahking notes (page 917), his inference concerning the exchange rate follows from the arbitrary exclusion of data at the beginning of the sample. An examination of Chart 1 reveals, however, that in excluding this part of the sample, Ahking is throwing away a disproportionate amount of the variability in the data, which will inevitably reduce the quality of the statistical inference.

Schwert (1987) suggests, on the basis of Monte Carlo evidence, that the Phillips-Perron tests may be biased towards rejecting the null hypothesis of non-stationarity much too frequently. As a cross-check, therefore, we also computed augmented Dickey-Fuller statistics based on higher-order autoregressions (as constructed by Ahking) as well as a unit root test statistic developed by Johansen (1988, 1990)--the maximal eigenvalue statistic--in each case allowing for a linear trend in mean. The results of applying these additional tests (Table 2) merely serve to confirm those already discussed: Ahking is probably mistaken in asserting that the exchange rate series is stationary about a time trend but Taylor and McMahon overlooked the presence of a deterministic trend in the U.K. wholesale price index. The U.K. price index <u>is</u> apparently stationary about a linear trend.

Does this state of affairs preclude the possibility of some form of long-run purchasing power parity holding? Not necessarily: on the evidence presented here, it appears that both the exchange rate series and the U.S. wholesale price series are realizations of unit root, I(1) processes while the U.K. price series is a realization from a process which is stationary about a linear trend. Thus, if the exchange rate and U.S. prices cointegrate, then this linear combination will be moving away from the U.K. price series only by a deterministic linear trend. Such a trend may be empirically small in magnitude (albeit statistically significant) and may be explicable economically, for example, in terms of capturing the effects of movements in variables such as relative productivity differentials or other structural changes during the sample period (Cassel, 1918; Balassa, 1964; Yeager, 1976).

Table 3 contains results of tests for cointegration between the exchange rate and U.S. prices using both the augmented Dickey-Fuller test applied to the cointegrating regression and a likelihood ratio (stochastic

Table 2. Dickey-Fuller and Johansen Unit Root Tests with Allowance for Trend in Mean

Series	Sample period	Lags in autoregressive representation	τ_τ	J_τ
Dollar-Sterling	1921:1-1925:5	3	-2.363	5.869
	1921:1-1925:5	3	-1.934	4.074
UK WPI	1921:2-1925:5	7	-5.097	20.183
	1921:2-1924:5	7	-4.370	19.314
US WPI	1921:1-1925:5	2	-3.096	9.507
	1921:1-1924:5	2	-2.490	6.353

* The null hypothesis is that there is a single unit root in the autoregressive representation. t is the augmented Dickey-Fuller statistic and J the Johansen maximal eigenvalue statistic (Johansen and Juselius, 1989, Johansen, 1990), each allowing for drift and trend in mean. Critical values are as follows (Fuller, 1976; Johansen and Juselius, 1989; Johansen 1990):

Critical values:	10 percent	5 percent	2.5 percent	1 percent
τ_τ :	-3.24	-3.60	-3.95	-4.38
J_τ :	6.691	8.083	9.658	11.576

Table 3. Tests for Cointegration of the Exchange Rate and U.S. Prices

$$e = a + bp$$

A) Ordinary least squares estimation

Sample period	$\hat{\alpha}$	$\hat{\beta}$	R^2	$\hat{\tau}_\mu$
1921:2-1925:5	-3.750	1.139	0.31	-4.205
1921:2 1924:5	-2.941	0.959	0.23	-4.480

B) Johansen estimation

Sample period	VAR lags	$\hat{\beta}$	$J(H_0:r\leq 1)$	$J(H_0:r=0)$	$LR(H_0:\beta=1)$
1921:4-1925:5	2	2.29	4.315	20.924	4.49 (0.03)
1921:5-1924:5	3	2.11	7.036	19.342	2.30 (0.13)

* R^2 is the coefficient of determination from the cointegrating regression; $\hat{\tau}_\mu$ denotes the augmented Dickey-Fuller statistic applied to the residuals from the cointegrating regression, with six lagged first differences in the auxiliary regression; critical values for this statistic are given below (Engle and Yoo, 1989). The Johansen (1988) maximum likelihood technique is reported only for the first (i.e., largest) eigenvalue of the stochastic matrix and after normalization on the exchange rate. The J statistics are likelihood ratio (stochastic matrix trace) statistics for the null hypothesis indicated in parenthesis where r denotes the number of cointegrating vectors, constructed as in Johansen, 1988; critical values for these statistics are given below (Johansen and Juselius, 1989; Johansen, 1990). LR denotes a likelihood ratio statistic for the null hypothesis indicated in parenthesis (i.e., β=1), constructed as in Johansen (1988) and is distributed as central chi-square under the null; figures given in parenthesis are marginal significance levels.

Critical values:	10 percent	5 percent	2.5 percent	1 percent
$\tilde{\tau}_\tau$:	-3.24	-3.60	-3.95	-4.38
$J(r\leq 1)$:	17.957	20.168	22.202	24.988
$J(r=0)$:	7.563	9.094	10.709	12.741

matrix trace) test developed by Johansen (1988). These tests, and estimates of the cointegrating parameter, were generated for both the full sample and for the sample excluding the last 12 data points. The augmented Dickey-Fuller statistics and the Johansen tests for the shorter sample size are in agreement--the hypothesis of non-cointegration is easily rejected at the 5 percent level. Using the Johansen method applied to the full sample, however, the likelihood ratio test for non-cointegration $[J(H_0:r=0)]$ is just insignificant at the 5 percent level, although significant at the 10 percent level. Since these critical values are only approximate, however, the general picture which emerges from these results is that the two series <u>are</u> cointegrated. 1/ The ordinary least squares estimates of the cointegrating parameter (which are 'super-consistent'--Stock, 1987) also suggest a value of the cointegrating parameter (normalized on the exchange rate) extremely close to unity. 2/ A formal test that the parameter is indeed unity, constructed as in Johansen 1988, does not reject the null hypothesis at the 5 percent significance level when the shorter sample is used but does so--albeit marginally--when the full sample is used. Again, we suggest that the weight of evidence suggests strongly that the two series are cointegrated with a cointegrating parameter equal to unity. In moving to the estimation of an error correction form, however, we deemed it prudent at this stage to impose cointegration and a unit cointegrating parameter for the shorter period only.

Table 4 contains an estimate of the parsimonious error correction form, in which U.K. prices are entered in log-levels and we have included a trend term in order to 'mop up' the deterministic trend in this series. 3/ The insignificant F-statistic for the exclusion of an additional first lag of the U.S. price series, FP^{US}, indicates that we are justified in imposing linear long-run homogeneity of the exchange rate with respect to the U.S. price level. The insignificant F-statistic for the exclusion of a quadratic trend term from the equation should dispel any worries that the U.K. series contains a quadratic trend (Ahking, page 917). Moreover, the equation is quite impressive in terms of the range of

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1/ Note that in testing for cointegration using two sample sizes and two methods, we are essentially testing the same hypothesis <u>four</u> times. This means that the true significance level of the <u>overall</u> test will probably be much greater than 5 percent--somewhere between 5 percent and 20 percent (=4 x 5 percent).

2/ An additional test that the parameter is unity will be given below when we discuss the estimated error correction form. Note that in testing this hypothesis twice with different sample sizes, the true significance level of the test of expanded--see footnote 1 above.

3/ Although the linear trend was insignificant on the basis of its t-ratio, its exclusion led to significantly worse diagnostics for the equation as a whole. Its inclusion is, in any case, consistent with the implications of the preceding analysis for the time series properties of the series under examination--see footnote 2 on page 12.

Table 4: Estimated Error Correction Form: 1921:5-1924:5 *

$$\Delta e_t = \underset{(0.105)}{-0.439}\,(e_{t-1}-p^{US}_{t-1}) + \underset{(0.134)}{0.419}\,\Delta e_{t-2} + \underset{(0.219)}{0.533}\,\Delta p^{US}_t + \underset{(0.227)}{0.333}\,\Delta p^{US}_{t-1}$$

$$\underset{(0.134)}{-0.573}\,p^{UK}_t + \underset{(0.121)}{0.279}\,p^{UK}_{t-3} + \underset{(0.00051)}{0.00050}\,t + \underset{(0.210)}{0.111}$$

$R^2 = 0.64$; SER =1.49%; DW = 1.70; $AR_{1-4}(4,25) = 0.88$; [0.49]

$AR_{5-6}(2,27) = 0.24$; [0.79] ARCH(4,21) = 0.04; [0.99] HET(14,14) = 0.76; [0.76]

RESET(1,28) = 1.29; [0.27] NORM(2) = 3.99; [0.14] CHOW(12,29) = 1.91; [0.08]

$Ft^2(1,28) = 0.007$; [0.93] $FP^{US}(1,28) = 1.59$. [0.22]

* R^2 is the coefficient of determination, SER the standard error of the regression and DW the Durbin-Watson statistic; AR_{i-j} is a Lagrange multiplier test statistic for serial correlation of order i to j; ARCH is a test statistic for up to fourth-order autoregressive conditional heteroscedasticity (Engle, 1982); HET is the Breusch-Pagan (1979) test for heteroscedasticity; RESET is Ramsey's (1969) test for functional misspecification; NORM is a test statistic for normality of the fitted residuals based on the coefficients of skewness and excess kurtosis; CHOW is Chow's (1960) test for predictive failure and parameter stability, obtained by using the estimated model to forecast 12 months forward out of sample up to May 1925; Ft and FP are test statistics for the exclusion of a quadratic trend term and one lag of the U.S. wholesale price index respectively. NORM is distributed as central chi-square with two degrees of freedom under the null hypothesis; all other statistics are central F with the indicated degrees of freedom. Figures in parenthesis denote estimated standard errors, those in brackets denote marginal significance levels.

diagnostic tests it passes and appears stable when used to forecast the exchange rate over the last 12 months of sterling's float (CHOW). 1/

IV. Long-Run PPP and the Norman Conquest of $4.86

Britain's decision to return to the Gold Standard at the prewar parity of $4.86 to the pound, largely instigated by the then Governor of the Bank of England, Montagu Norman, was seen by Keynes (e.g., 1925) as an overvaluation of sterling of at least 10 percent--an opinion with which many commentators since have concurred (e.g., Ashworth, 1960; Pollard, 1962; Kindleburger, 1964) and which seems to be the received wisdom on the issue (Moggridge, 1972). There are, however, informed commentaries--both contemporary (e.g., Cassel, 1926; Gregory, 1926 2/) and more recent (Walter, 1951; Morgan, 1952; Youngsen, 1960)--which suggest that sterling either was <u>undervalued</u> against the dollar at $4.86 or at least was not overvalued quite to the degree suggested by Keynes and others. The above analysis can, in fact, be used to shed some light on this heated issue.

1/ Estimating the error correction representation over the period up to and including May 1925 yields the following results:

$$\Delta e_t = \underset{(0.089)}{-0.369}(e_{t-1}-p^{US}_{t-1}) + \underset{(0.123)}{0.362}\,\Delta e_{t-2} + \underset{(0.191)}{0.690}\,\Delta^{US}_t + \underset{(0.175)}{0.362}\,\Delta p^{US}_{t-1}$$

$$- \underset{(0.108)}{0.493}\,p^{UK}_t + \underset{(0.080)}{0.284}\,p^{UK}_{t-3} + \underset{(0.00022)}{0.00077}\,t - \underset{(0.210)}{0.105}$$

$R^2 = 0.53$; SER = 1.52%; DW = 1.61; $AR_{1-4}(4,37) = 1.25$; [0.31]

$AR_{5-6}(2,39) = 0.56$ [0.58]; $ARCH(4,33) = 0.52$ [0.72]; $HET(14,26) = 1.00$ [0.48];

$RESET(1,40) = 0.25$ [0.62]; $NORM(2) = 1.97$ [0.37]; $Ft^2(1,40) = 0.100$ [0.75]; $FP^{US}(1,28) = 1.587$ [0.22]

2/ Although see also, Gregory (1957, 1968).

Solving the error correction form reported in Table 4 for its long-run steady-state solution 1/ yields the following expression for the equilibrium exchange rate:

$$\text{Equilibrium rate} = \exp\left\{p_t^{US} - 0.670\ p_t^{UK} + 0.253 + 0.00114t - 0.0169\ \pi^{UK}\right\} \quad (4)$$

where π^{UK} is the percentage steady-state monthly rate of U.K. wholesale price inflation. 2/

Although the equilibrium exchange rate, as expressed in (4), displays asymmetric responses to movements in British and American prices, this finding is in common with analyses of, for example, dollar-sterling during the 1930s and may be explicable in terms of differing levels of trade protection or other market distortions (see, e.g., Cassel, 1918; Broadberry, 1987; Taylor, 1988; Broadberry and Taylor, 1990). 3/

Substituting an estimate of π^{UK} into (4) 4/, together with the values of p_t^{US}, p_t^{UK} and the trend term for May 1925 yields an estimated equilibrium value for sterling against the dollar of $4.63 per pound sterling. This suggests an overvaluation of sterling against the U.S. dollar on sterling's return to the Gold Standard only of the order of some 5 percent.

Using equation (4), we have graphed in Chart 2 the actual and equilibrium dollar-sterling exchange rate over the whole period and for 12 months after the return to gold. Examination of the two series suggests that movements in relative prices had moved the equilibrium rate close to the imposed parity by the end of 1925 and that, by mid-1926, sterling was in

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1/ The steady-state rate of U.S. inflation is set to zero in solving for the long-run solution. This seems empirically justified here since, estimating a first-order autoregression for this series (i.e., Δp_t) over the whole sample period and solving for the steady-state yields an estimate of long-run monthly inflation of just 0.0175 percent (an annual rate of 0.21 percent) with an asymptotic t-ratio of 0.052 (Bardsen, 1988). Similar results were obtained with autoregressions of order three and six. On the other hand, solving a (sixth-order) autoregression in the log-level of U.K. prices, including a trend term, yielded a long-run monthly rate of inflation of 0.2 percent (an annual rate of 2.4 percent).

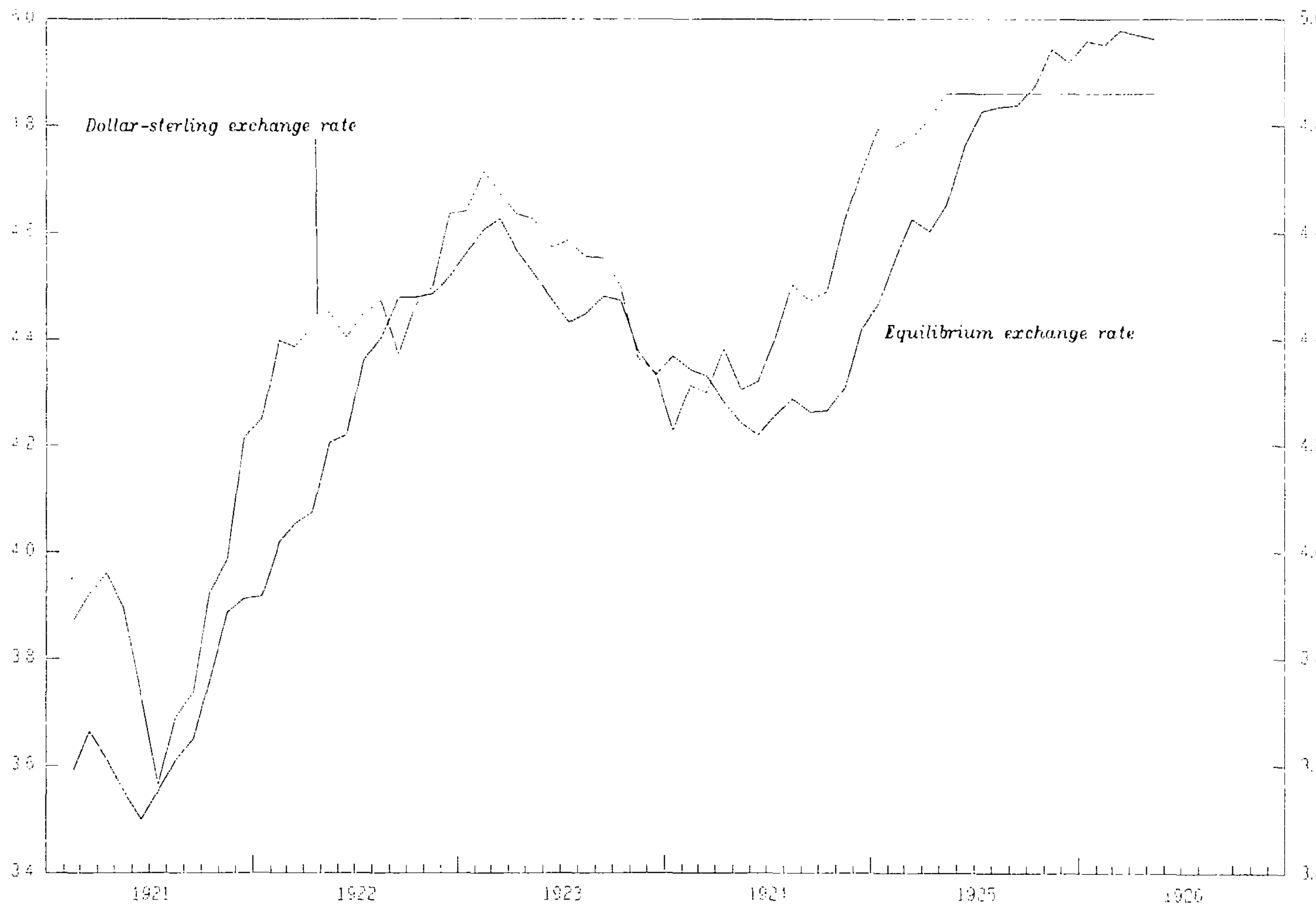
2/ Note that, with a long-term trend coefficient of 0.002 in the U.K. wholesale price series (see previous footnote), the trend in the equilibrium level of the exchange rate has a coefficient of $(0.0014 - 0.67 \times 0.002) \simeq 0$; which is consistent with our finding of no deterministic trend in the exchange rate series.

3/ Imposing a long-run unit coefficient on U.K. prices in the error correction representation yields an F-statistic of $F(1,29) = 5.46$, which has a marginal significance level of 2.65 percent.

4/ See footnote 1 above.

CHART 2

US-UK EXCHANGE RATE AND EQUILIBRIUM EXCHANGE RATE 1921-26



fact undervalued against the dollar by some 2 percent. 1/ Although these figures may seem to fly in the face of the received view, they are not in fact widely different from those reported in other recent studies (e.g., Redmond, 1984), as well as the earlier studies cited above. Moreover, it must be remembered that this analysis is entirely in terms of the bilateral dollar-sterling rate. Redmond's (1984) analysis strongly suggests that sterling was in fact heavily overvalued at the prewar parity against a number of other (particularly European) currencies 2/ and that the true economic measure of overvaluation should relate to a trade-weighted exchange rate index.

V. Conclusion

The value of Ahking's contribution lies in drawing attention to the presence of a deterministic trend in the U.K. wholesale price index series during the period under investigation. By carefully incorporating this into an econometric analysis, however, the main implications of the Taylor-McMahon paper are, if anything, amplified: some form of long-run purchasing power parity appears to have held between the United States and the United Kingdom for virtually the whole of the 1920s float.

1/ It is also apparent from Chart 2 that sterling had become increasingly overvalued from early 1924 until mid-1925, presumably due to speculative inflows (Aliber, 1962). The chart reveals, however, that the drift away from the underlying fundamentals may not have been as dramatic or as permanent as originally suggested by Taylor and McMahon.

2/ See also, Walter (1951), pages 14-17.

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