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US Dollar / British Pound Market

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EXPECTATIONS AND FORECASTING IN THE US DOLLAR/BRITISH POUND MARKET

by

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EXPECTATIONS AND FORECASTING IN THE US DOLLAR/BRITISH POUND MARKET

ABSTRACT

Critics have pointed to the poor performance of traditional economic models in forecasting exchange rates, and have drawn attention to perceived shortcomings of these models, such as undue reliance on single equation methods, and inadequate representation of expectations. This paper addresses these issues. A simultaneous, rational expectations model of exchange rate determination is presented and estimated, using information from both spot and futures markets. The model contains separate functional relationships for short and long hedgers, short and long speculators, as well as a spot rate equation.

Augmented Dickey-Fuller and Phillips-Perron tests for unit roots suggest that spot and futures exchange rates and interest rates are non-stationary, although results for exports and imports are ambiguous. Engle-Granger cointegration tests suggest that these I(1) variables are cointegrated, and the Johansen procedure indicates that there is one cointegrating relationship in each of the respective equations. The errors of most of the structural equations exhibit low order ARCH effects; an exception is the short speculation function, which exhibits first order serial correlation. All estimates of structural parameters have the anticipated signs, and all are significant. Post-sample, this model forecasts spot and futures exchange rates with per cent RMSE's less than 2%, which clearly outperforms forecasts by a random walk model.

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EXPECTATIONS AND FORECASTING IN THE US DOLLAR/BRITISH POUND MARKET

I

INTRODUCTION

Critics such as Meese (1990), Isard (1987) and Wolff (1987) have claimed that exchange rate models produced by economists during the past two decades have performed poorly, in the sense that such models explain only a small proportion of exchange rate variation, and they cannot surpass a random walk model in post-sample forecasting. For example, Meese (1990, p.117) argued that the "... proportion of .... exchange rate changes that current models can explain is essentially zero", while Isard (1987, p.3) thought that these models "explain little of the observed variances of exchange rates during the 1970s and 1980s." Inadequacies of exchange rate models to which these authors drew attention include insufficient allowance for simultaneity (Meese, 1990, p.117) and the use of expectations hypotheses not fully consistent with the structural models employed (Isard, 1987, p.16).

The objective of this paper is to develop and estimate a model of exchange rate determination, using information from both spot and futures markets. In particular, functional relationships will be developed for short hedgers and short speculators in futures, for long hedgers and long speculators in futures, as well as a spot rate equation. This model, which has its foundations in the theoretical work of Peston and Yamey (1960), extends the empirical work of Giles *et al.* (1985), Goss (1990), and Goss *et al.* (1992) on storable, consumable commodities, to foreign exchange markets, and extends the work of Goss and Avsar (1996), on foreign exchange, to more active markets.

Expectations are represented according to the rational expectations hypothesis, and the model is estimated by instrumental variables. The performance of the model is evaluated *inter alia* by hypothesis tests on the various parameter estimates, and by assessing the intra- and post-sample forecasting ability of the model. This latter exercise includes a comparison with a random walk model, and with a lagged futures rate which permits a test of the efficient markets hypothesis.

The US Dollar/British Pound (US$/BP) market has been studied by many, including Hansen and Hodrick (1980), Baillie, Lippens and McMahon (1983), and Boothe and Longworth (1986), all of whom rejected the unbiasedness hypothesis with different data sets, although, because this hypothesis enjoins the hypotheses of rational expectations and risk neutrality, it is not clear which component of the joint test was rejected. Rejection of unbiasedness has been interpreted by some authors (e.g. Hodrick and Srivastava, 1984, 1987) to indicate the presence of a time varying risk premium, in the presence of rational expectations, and by others (e.g. Bilson, 1981; Taylor, 1992) as evidence of market inefficiency (see the survey by Hodrick, 1987, especially pp. 54-56 and pp. 140-50).

The specification of the model is discussed in section II of this paper, while section III discusses the data and estimation methods employed, including tests for unit roots and cointegration. Intra-sample results are presented and evaluated in section IV, while post-sample forecasts are discussed in section V. Some conclusions are presented in section VI.

II

SPECIFICATION OF THE MODEL

This section discusses the specification of the functional relationships for the various categories of economic agent mentioned above. Consider first the position of short hedgers.

These agents, such as US exporters to UK or British investors undertaking capital inflow to the US, have long spot market commitments, and are hedging the risk of a fall in the spot rate. Two of the main types of hedging distinguished by Working (1953, 1962) are carrying charge hedging and selective hedging. On the hypothesis that short hedgers in this market are of the carrying charge type, their commitments in the US$/BP futures market could be expected to vary directly with the current forward premium (i.e. futures rate less spot rate) and negatively with the expected forward premium. If, on the other hand, short hedgers in this market are selective hedgers (where a proportion only of their spot commitments is hedged), their futures market commitments would be expected to vary directly with the current futures rate and negatively with the expected futures rate. Preliminary estimation favoured the latter of these two hypotheses.

The futures market commitments of short hedgers can be expected also to vary directly with a measure of their spot market commitments, such as US exports to UK, and/or British capital inflow to US. Alternatively, the volume of short hedgers futures market commitments might be expected to vary directly with the magnitude of net long speculation in this market, as discussed by Keynes (1930, pp. 142-44). Preliminary estimation favoured the former of these two alternatives. Hence the final specification for this equation is

$$SH_t = \theta_1 + \theta_2 P_t + \theta_3 P^*_{t+1} + \theta_4 X_t + \theta_5 BKI_t + e_{1t} \tag{1}$$

where SH = futures market commitments of short hedgers (see data section below for measurement of market commitments);

P = current futures rate;

P^*_{t+1} = rational expectation of the futures rate at time (t + 1) formed at time t;

X = US exports to UK;

BKI = British capital inflow to USA;

t = time in months;

e = error term;

and θ_1 = constant, θ_2, θ_4, $\theta_5 > 0$, $\theta_3 < 0$.

Both the futures rate (P) and the spot rate (A) are measured as US$/BP, so that a rise in price means devaluation of the US currency.

The rational expectations hypothesis (REH), which represents expectations in this model, originated with Muth's observation that mean expectations in an industry are as accurate as 'elaborate equation systems' and his suggestion that 'rational' expectations are the same as the predictions of the relevant economic theory (Muth 1961, p. 316).

The REH assumes first, that agents efficiently use all relevant publicly available information at time t in forming their expectations about an economic variable at time t + 1. This implies that expectational errors will have an expected value of zero, and will be uncorrelated with elements of the information set, and with past expectational errors. Second, the REH assumes that agents correctly anticipate future prices, in the sense that the subjective probability distributions of agents coincide with the objective probability distribution of the system. Third, the REH implies that agents have this particular economic model in mind in forming their expectations, so that any test of the REH is a joint test of the expectations hypothesis and the economic model in question (Maddock and Carter, 1982).

The question is then how do agents learn to form rational expectations. Blume, Bray and Easley (1982) suggest that this outcome can be achieved by agents using the same forecasting rule consistently over a long period, while Hirsch and Lovell (1969) emphasize having a long history in the industry in question. To the question of the likelihood of agents learning to form rational expectations, the answer of Bray and Savin (1986) is more optimistic than that of Frydman (1983, pp. 106-117). Experimental evidence to support the view that markets converge to a rational expectations equilibrium in a comparatively short period of time, as required by this model, can be found in Plott and Sunder (1982), Friedman, Harrison and Salmon (1983) and Harrison (1992). Moreover, the evidence in these last two papers suggests that the convergence is more rapid with futures markets than without.

Long hedgers are agents with forward actuals commitments such as US importers or US investors exporting capital to the UK, and they are hedging the risk of a rise in the spot exchange rate. Traditionally the futures market position of the long hedger has been regarded as the mirror image of the short hedger's position (Stein 1961), although a qualification to the accuracy of this description has been noted (Yamey 1971). On the 'carrying charge' hedging hypothesis, the market commitments of long hedgers would be expected to vary negatively with the current forward premium and directly with the expected forward premium. On the selective hedging hypothesis, the futures market positions of long hedgers would be expected to vary negatively with the current futures rate and directly with the expected futures rate. Preliminary estimation supported the latter of these two hypotheses.

The market commitments of long hedgers could be expected also to vary directly with a measure of their forward actuals commitments, such as US imports from UK, and US capital outflow to Britain; alternatively, long hedgers' market commitments could be expected to vary directly with the volume of net short speculation. Preliminary estimation favoured the former of these two alternatives.

Hence the specification of this equation is

$$LH_t = \theta_6 + \theta_7 P_t - \theta_8 P^*_{t+1} + \theta_9 M_t + \theta_{10} AKO_t + e_{2t} \qquad (2)$$

where LH = futures market commitments of long hedgers;

M = US imports from UK;

AKO = US capital outflow to Britain;

and $\theta_7 < 0$; θ_8, θ_9 $\theta_{10} > 0$.

The specifications of (1) and (2) are consistent with the view that short and long hedgers pursue the dual objectives of risk reduction and expected gain (see Stein 1961, and Johnson 1960).

Short speculators in futures expect the futures exchange rate to fall, and can be expected to vary their market commitments directly with the current futures rate, and negatively with the expected futures rate. Traditionally, the market commitments of speculators have been expected to vary negatively also with the marginal risk premium, if these agents are assumed to be risk averse (see Kaldor 1953, Brennan 1958). This traditional view of the influence of the risk premium has been challenged recently by Stein (1986, pp.48-52) who argues in his "hedging pressure theory" that an increase in the risk premium may have a positive or negative influence on the exchange rate, and hence by implication, on the market commitments of short speculators. Market commitments of short speculators could be expected to vary positively also with the magnitude of net long hedging, extending the ideas of Keynes (1930, pp. 142-44) to short speculation.

While the US-UK interest differential could be expected to have a positive impact on the commitments of short speculators, preliminary estimation did not support the inclusion of this variable. Hence the specification of the short speculation in futures equation is

$$SS_t = \theta_{11} + \theta_{12} P_t + \theta_{13} P^*_{t+1} + \theta_{14} r_t + \theta_{15} NLH_t + e_{3t} \quad (3)$$

where SS = futures market commitments of short speculators;

r = marginal risk premium;

NLH = net long hedging

$\equiv$ LH - SH;

and $\theta_{12}, \theta_{15} > 0; \theta_{13} < 0; \theta_{14} \gtrless 0.$

Long speculators in futures, on the other hand, expect the futures rate to rise, and buy British Pound futures in support of their expectations. Their market commitments, therefore, could be expected to vary negatively with the current futures rate, directly with the expected futures rate, and in the traditional view (e.g. Kaldor, 1953) negatively with the marginal risk premium, although the argument of Stein (1986, pp.48-52) regarding the ambiguity of the effect of the risk premium applies equally to long speculators in futures. The US - UK interest differential could be expected to have a negative impact on the market commitments of these agents, although preliminary estimation did not support the inclusion of this variable. The market commitments of long speculators could be expected to vary directly also with the magnitude of net short hedging, on the argument of Keynes (1930, pp. 142-44). Hence the specification of this equation is

$$LS_t = \theta_{16} + \theta_{17} P_t + \theta_{18} P^*_{t+1} + \theta_{19} r_t + \theta_{20} NSH_t + e_{4t} \qquad (4)$$

where LS = market commitments of long speculators in futures;

NSH = net short hedging

$\equiv$ SH - LH;

and $\theta_{17} < 0$; θ_{18} , $\theta_{20} > 0$; $\theta_{19} \gtrless 0$.

The final relationship in this model is the spot rate equation, in which the spot rate is hypothesized to vary directly with the futures rate and US exports to UK, and negatively with the US interest rate; preliminary estimation indicated that imports, capital flows and the interest differential should not be included. Hence the specification of this function is:

$$A_t = \theta_{21} + \theta_{22} P_t + \theta_{23} X_t + \theta_{24} IA_t + e_{5t} \qquad (5)$$

where IA = US interest rate;

and θ_{22} , $\theta_{23} > 0$; $\theta_{24} < 0$.

The model contains six endogenous variables: SH, LH, SS, LS, P, A, and is completed with the following identity:

$$SH + SS \equiv LH + LS \qquad (6)$$

where (6) is a futures market clearing identity.

Conventional identification conditions do not apply to linear multi-equation models with forward rational expectations (Pesaran 1987, p.119). The model developed here, however, fulfils the conditions developed by Pesaran (1987, p.156-60).

III

DATA, UNIT ROOTS, COINTEGRATION TESTS AND ESTIMATION

DATA

This section discusses the data definitions and sources under the headings `Endogenous Variables' and `Exogenous Variables'.

Endogenous Variables

Commitments of traders for the variables SH, LH, SS, LS are open positions at end of month measured in number of contracts and converted to million British pounds.[1] These data are collected by the Commodity Futures Trading Commission (CFTC) in the USA for `Reporting' (Large) `Commercials' (Hedgers) and `Non-commercials' (Speculators), and for `Non-reporting traders', for both long and short positions. Data for `non-reporting (small) traders' are not classified by the CFTC as between hedging and speculation, and in some studies it has been suggested that for certain commodities for some time periods these data should be treated as all speculative (e.g. see Peck (1982)), while in other studies the open positions of small traders have been divided between hedging and speculation in the same ratio as the open positions of reporting traders (e.g. see Goss *et al.* (1992)). After experimentation, it was decided in this study to include all open positions for non-reporting traders as speculation. These data were not collected by the CFTC from 1981(12) to 1982(12), and this has implications for the definition of the sample period (see below).

Data for the spot rate (A) are daily observations of the bid side, last quote (interbank rate) on the median trading day of each month, quoted in US dollars per British pound, from the IMM *Yearbook*, 1978-88.

Data for the futures rate (P) are daily observations of the last trade on the median trading day of the month, quoted in US dollars per British pound, for a futures contract which is on average two months prior to maturity.[2] These quotations are taken from the IMM *Yearbook*, 1978-88.

Exogenous Variables

Data on exports (X) are monthly observations on US exports to Britain from the *Survey of Current Business*, converted to million British pounds. Data on imports (M) are monthly observations on US imports from Britain also from the *Survey of Current Business*, converted to million British pounds.

Interest rate observations (IA) are US prime rate monthly averages in per cent per annum from the *Federal Reserve Bulletin*. Marginal risk premium data (r) are monthly averages of the US 90 day Commercial Paper Rate per cent per annum, less the monthly average of the US 90 day Treasury Bill rate per cent per annum, both these rates being taken from the *Federal Reserve Bulletin*.[3]

Data on US capital outflows to Britain (AKO), and on capital inflow from UK to US (BKI), are generated from changes in UK liabilities (claims) of banks and other institutions, both public and private sector, to (on) US at end of each quarter (i.e. 31 Mar., 30 Jun., 30 Sept., 31 Dec.) in mill. British pounds from the Bank of England *Quarterly Bulletin*.[4] Data on AKO and BKI were interpolated to monthly observations with the program TRANSF (Wymer 1977).

The basic sample period is from 1978(10) to 1986(12). As stated above, no CFTC open position data were collected from 1981(12) to 1982(12), so that after allowance for lags and leads, the effective sample period is from 1978(10) to 1981(11), and from 1983(04) to 1986(11), which is a total of 82 observations. The post-sample period dates from 1987(01) to 1988(12), comprising 24 observations.

Unit Roots and Cointegration Tests

To obtain meaningful estimates of the parameters of the model, it is necessary that the residuals of the estimating equations are stationary. This condition will be fulfilled if all the variables in these equations are stationary (i.e. integrated of order I(0)), or alternatively, if some of these variables are integrated of order I(1), this condition will be fulfilled only if the I(1) variables are cointegrated.

Equation (7) is an autoregressive representation of a time series in Z, an economic variable

$$Z_t = \rho Z_{t-1} + e_t \tag{7}$$

where ρ is a real number, and e_t is NID$(0,\sigma^2)$. If $|\rho| < 1$, Z_t converges to a stationary series as $t \rightarrow \infty$, while if $\rho = 1$ there is a single unit root, and Z_t is non-stationary; (if $|\rho| > 1$, the series is explosive). Tests of the hypothesis $H(\rho=1)$ for the model in (7), and for variations of this model with constant and time trend, were developed by Dickey and Fuller (1979, 1981). These tests were extended by Said and Dickey (1984) to accommodate autoregressive processes in e_t of higher but unknown order. In this latter case the model is augmented by lagged first differences in Z to whiten e_t, and the hypothesis $H(\rho=1)$ is tested by the Augmented Dickey-Fuller Test (ADF).[5]

Unit roots in exchange rates have been studied by several authors and almost invariably it has been found that the hypothesis of a single unit root cannot be rejected (see the surveys in Hodrick (1987, p.29) and Baillie and McMahon (1989, pp.106-109)). In this paper the following models were estimated by ordinary least squares (OLS) to test the hypothesis of a unit root in all endogenous and exogenous variables in the model:

$$\Delta Z_t = \mu + \beta t + \gamma Z_{t-1} + e_t \tag{8}$$

$$\Delta Z_t = \mu + \beta t + \gamma Z_{t-1} + \phi \Delta Z_{t-1} + e_t \tag{9}$$

$$\Delta Z_t = \mu + \gamma Z_{t-1} + \phi \Delta Z_{t-1} + e_t \tag{10}$$

where $\gamma = \rho - 1$;

μ = constant;

β, ϕ are coefficients to be estimated;

e_t is assumed to be NID$(0,\sigma^2)$.

The hypothesis H(ρ=1) is addressed by testing the hypothesis H(γ=0) in (8)-(10). This is executed by the ADF test, although it is preferable to refer to critical values of Mackinnon (1991), which are based on more replications than the original Dickey-Fuller tables. Estimates of $\hat{\gamma}$, calculated ADF test statistics and 5 per cent critical values from Mackinnon (1991) for each variable in the model, are given in Appendix 1, the choice of model depending upon the lag length necessary to whiten e_t. These tests support the view that the following variables are integrated of order I(0): SH, LH, SS, LS, AKO, BKI, r, NLH, NSH, while the following variables are I(1) : P, A, X, M and IA. (The instruments for P^*_{t+1}, and A^*_{t+1} are also I(1); see below ESTIMATION for an explanation of the formation of these instruments).

To address the issue of higher order autocorrelation in (7), the method of Phillips and Perron (1988) makes a non-parametric correction to the estimated test statistic, to allow for the autocorrelation which would otherwise be present in the residuals. Asymptotically, the same limiting distributions apply as in the Dickey-Fuller case, and the same critical values may be employed. Phillips-Perron tests for unit roots were conducted for all variables in this model (not reported here for reasons of space), with the same indications about stationarity as under the Dickey-Fuller tests, except for the export (X) and import (M) variables, where the unit root hypothesis was rejected under the Phillips-Perron tests.

The Phillips-Perron procedure, however, appears to suffer greater size distortion than the Said-Dickey procedure (i.e. rejects a true hypothesis a proportion of the time greater than the nominal size), especially with negative moving average errors. (See Banerjee *et al.* 1993, pp. 108-109, 113, 129.) This point tends to favour retention of the decision, made under the ADF tests, to regard X and M as I(1). In any case, as will be seen below, the $\bar{X}$ and M variables appear to be cointegrated with other I(1) variables in the equations in which they appear, and so the specification implications of both sets of unit root tests are the same.

As stated above, for the residuals of the structural equations to be stationary, it is necessary for the I(1) variables in these equations to be cointegrated. To investigate this question, the cointegration test analysed by Mackinnon (1991), which is based on the work of Engle and Granger (1987), was employed. This test requires first that a relationship between the I(1) variables (the cointegrating equation) be estimated by OLS. In the case of equation (1) this gives rise to

$$P_t = \alpha_0 + \alpha_1 \hat{P}_{t+1} + \alpha_2 X_t + u_t \tag{11}$$

The residuals from this process can be written as

$$\hat{u}_t = P_t - \hat{\alpha}_0 - \hat{\alpha}_1 \hat{P}_{t+1} - \hat{\alpha}_2 X_t \tag{12}$$

where $\hat{u}_t$ represents estimated values of the residuals, and $\hat{\alpha}_0$, ..., $\hat{\alpha}_2$ refers to estimated values of the parameters of (11). The hypothesis of no cointegration in (11) is addressed by testing the hypothesis that $\hat{u}_t$ in (12) has a unit root. If this latter hypothesis is rejected, then the hypothesis of no cointegration in (11) is rejected.

To test the hypothesis of a unit root in $\hat{u}_t$ the following model was estimated:

$$\Delta \hat{u}_t = \mu + \gamma \hat{u}_{t-1} \tag{13}$$

and the hypothesis H(γ=0) was tested. As the information in Appendix 2 shows, this hypothesis can be rejected at the one per cent level, and hence the hypothesis of no cointegration in (11) is rejected. This outcome supports the view that the residuals in (1) are stationary. The same procedure was employed for equations (2) to (5), and, as the test statistics in Appendix 2 show, in all cases the hypothesis of a unit root, in the residuals of the cointegrating equation, can be rejected. These results support the view that the residuals of equations (1) to (5) are stationary; (in the case of equation (5) only, it was necessary to include the term $\phi\Delta\hat{u}_{t-1}$ to whiten the error term). As a consequence of these cointegration tests no re-specification of the structural model is necessary.

The Engle-Granger procedure has been criticized *inter alia* on the grounds first, that the distribution of the test statistics is not independent of the nuisance parameters of the particular application, and second, that it is capable of estimating one cointegrating vector only (which varies according to the normalization). The procedure of Johansen (1988) and Johansen and Juselius (1990) overcomes these difficulties, and their likelihood ratio test is capable of identifying all cointegrating vectors in a set of I(1) variables. Two tests have been developed by these authors: the first is the "trace" test, which tests the hypothesis that the number of cointegrating vectors m is at most equal to q (where q < n, the number of I(1) variables in the relationship), against the general alternative that m ≤ n. The second test, the "λ max" test, tests the hypothesis that m ≤ q, against the specific alternative m ≤ q + 1.

In this paper, both these tests have been employed to investigate the number of cointegrating vectors in equations (1), (2) and (5) (equations containing more than two I(1) variables). Appendix 3 reports the likelihood ratio test statistics for the λ max test, together with 5 per cent critical values, for various values of m. It will be seen that in each case the hypothesis m = 0 is rejected, and the hypothesis m ≤ 1 is not rejected, suggesting that there is one cointegrating vector in each of these equations. The results of the trace test (not reported here for reasons of space) indicate the same outcomes, although this need not necessarily be the case (see Francis and Leachman, 1994, pp. 44-48).

The unit root and cointegration tests discussed in this section were executed with the package E Views - Micro TSP (Hall, Lilien and Johnston, 1994).

ESTIMATION

Full information estimators for simultaneous models with forward rational expectations are less robust to specification errors, and are computationally more demanding than limited information methods (Pesaran 1987, p.162). For these reasons the model presented here is estimated by the instrumental variable (IV) method of McCallum (1979), modified for recent developments in econometrics. This requires first, that an instrument is obtained, by OLS, for the expectation of an endogenous variable, as a fitted value on an information set comprising all exogenous and predetermined variables in the model. If the residuals of the structural equations are not serially correlated, those equations can be estimated by IV and this method will result in consistent estimates. This procedure is discussed by McCallum (1979) and Cumby *et al.* (1983) and is summarized in Giles *et al.* (1985, pp.754-55).

When serial correlation is present, however, a simple autoregressive (AR) correction with IV estimation will not produce consistent estimates, as Flood and Garber (1980) pointed out. In this case an AR transformation has been made, and each of the variables in the transformed equation was regressed on the elements of the relevant information set, using OLS. The fitted values so obtained were substituted in the transformed equation (see McCallum (1979, pp.67-68)), and consistent estimates of the parameters in that equation can be obtained by non-linear least squares, using the option LSQ in TSP (Hall, *et al.* 1993). This procedure was employed for equation (3).

The variance of the residuals of regression relationships for financial time series typically varies over time and is autocorrelated. This effect has been represented by the autoregressive conditional heteroscedasticity (ARCH) model (see Engle 1982, 1983). If e_t, the error term in a regression equation, is assumed to be N(0, σ^2), an ARCH (p) process postulates that, conditional on information at time (t-1)

$$\sigma_t^2 = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 e_{t-2}^2 + \dots \alpha_p e_{t-p}^2$$

Tests were executed for the presence of ARCH effects in equations (1) to (5), and, on this basis, the error structure in (1) was represented as an ARCH (1) process, while the error terms in (2) and (5) were each modelled as an ARCH (2) process, and the equations were re-estimated by maximum likelihood (ML). Similarly, the error structure in (4) was represented as an MA(1) process, and (4) was re-estimated by ML. No significant ARCH effects were evident in equation (3).

IV

RESULTS: INTRA-SAMPLE PERIOD

The parameter estimates of equations (1) to (5), together with asymptotic t values, are presented in Table 1. It will be seen, first, that the estimates of all nineteen structural parameters have the expected sign, and all are significant at the five per cent level (one tail test). These results provide strong support for the model specification developed above, and in view of the significance of the estimates of θ_3, θ_8, θ_{13}, θ_{18}, these results provide support for the rational expectations hypothesis. Second, estimates of the coefficients of the marginal risk premium, θ_{14} and θ_{19} in equations (3) and (4) respectively, are both negative and significant, thus lending support for the conventional view of the marginal risk premium, as developed by Kaldor (1953), Brennan (1958) and Telser (1958).

Third, the estimates of the coefficients of the two capital flow variables, BKI and AKO, employed here as measures of the spot market commitments (in part) of short and long hedgers respectively, are both significant, thus supporting the suggestion of Isard (1987) that capital flow variables have a role to play in modelling exchange rate determination. It should be noted that in the functional relationships for both short and long hedgers, equations (1) and (2), the results support the view that hedgers in both categories are of the "selective" rather than the "carrying charge" type, and are likely to hedge a fraction other than unity, of their spot market commitments.

Fourth, the estimates of the coefficients of NLH and NSH, in equations (3) and (4) respectively, are both significant, thus supporting the view of Keynes (1930), that the commitments of speculators in futures tend to vary negatively with the net commitments of hedgers on the opposite side of the market. Finally, it should be noted that the estimated coefficients of the ARCH, AR and MA processes, representing the error structures of the individual equations, are all significant, except the estimate of the ARCH(1) coefficient in equation (5), where the error term is represented by an ARCH(2) process.

An important test of the performance of a model such as this is its ability to forecast spot and futures exchange rates within the sample period, according to selected criteria. Table 2 provides an evaluation of intra-sample (static) simulation according to correlation coefficient, Theil's inequality coefficient and per cent root mean square error of forecast.[6] Concentrating on the per cent RMSE criterion, it can be seen that the model forecasts both spot and futures rates with per cent RMSE's of 2.8% which is comparable with the intra-sample performance of a similar model of the Australian dollar/US dollar exchange rate (see Goss and Avsar, 1996).

V

POST-SAMPLE SIMULATION

An evaluation of post-sample (dynamic) simulation of spot and futures exchange rates in this market, two months ahead,[7] is provided in Table 3, for the period 1987(01) to 1988(12) (24 observations). Concentrating again on the per cent RMSE criterion, prediction of both spot and futures rates has improved compared with intra-sample simulation of these rates, and in each case the per cent RMSE is less than two per cent. When compared with other predictors of the spot rate (see Table 4) it can be seen that this model (AS) outperforms the lagged futures rate (P_{t-2}) and the random walk model (RWALK2) according to the same criterion.[8] The first of these comparisons suggests that this model provides evidence against the semi-strong efficient markets hypothesis, because the model evidently contains information which is not reflected in the futures price. The EMH should not be rejected, however, until there is evidence that a model such as this can yield significant profits in excess of transaction costs (a sufficient condition for market inefficiency; see Leuthold 1991, pp. 66-70).[9] The second comparison suggests that the post-sample forecasting performance of this model is clearly superior to that of a random walk model, the latter being a conventional benchmark in exchange rate forecasts. According to critics, the forecasting performance of structural models of exchange rates is usually inferior to that of a random walk (see Section I above), so that this comparison provides some justification for the approach adopted here.

VI

CONCLUSIONS

This paper develops and presents estimates of a simultaneous, rational expectations model of exchange rate determination in the US dollar/British pound market, using information from both spot and futures markets. The model contains functional relationships for short hedgers, long hedgers, short speculators and long speculators, as well as a spot rate equation and market clearing identity. Estimation is by non-linear least squares in the presence of first order autocorrelation, and by maximum likelihood when there is evidence of ARCH effects. Model performance is evaluated *inter alia* by comparing post-sample predictions of the spot rate, produced by the model, with forecasts derived from a lagged futures rate and from a random walk model. The paper not only provides an analysis of exchange rate determination, but addresses also concerns raised by critics who have drawn attention to the poor forecasting performance of traditional exchange rate models, and who emphasized the perceived limitations of traditional approaches, such as undue reliance on single equation methods, inadequate representation of expectations, and insufficient attention to capital flows.

The main conclusions are as follows. First, the estimates of all nineteen structural parameters have the anticipated sign, and all are significant at the five per cent level, thus providing support for the model specification employed here, and for the rational expectations hypothesis. Second, while the results support the view that both short and long hedgers pursue the dual objectives of risk reduction and gain, these hedgers would appear to be of the "selective" rather than the "carrying charge" type, in the sense of Working (1953), thus hedging a proportion other than unity, of their spot market commitments. Third, capital inflow from Britain to US, and US capital outflow to Britain, employed here as (partial) measures of the spot market commitments of short and long hedgers respectively, appear to have a significant role to play in determining the market commitments by these hedgers, and hence in exchange rate determination.

Fourth, the market commitments of both short and long speculators respond not only to current and expected futures exchange rates, as economic theory suggests, but vary directly with the net commitments of hedgers on the opposite side of the market, as suggested by

Keynes (1930). Fifth, there is evidence of low order ARCH effects in three of the five structural equations in this model. The error term in one of the remaining equations is represented by a first order AR process, and in the other as an MA1 process.

Finally, in post-sample simulation, a model derived forecast of the spot rate clearly outperforms forecasts provided by, first, a lagged futures rate, and second, by a random walk model. The first comparison provides evidence against the semi-strong EMH, although this hypothesis should not be rejected until there is evidence that such a model can be used to produce significant profits net of transaction costs. The second comparison shows that the model clearly outperforms a conventional benchmark in exchange rate forecasting, and hence provides further justification for the approach adopted in this paper.

TABLE 1

PARAMETER ESTIMATES

Equation	Coefficient	Variable	Estimate	Asymp. t Value
(1)	θ_1	Const.	348.579	1.995
	θ_2	P_t	1040.83	3.633
	θ_3	P^*_{t+1}	-1327.28	-4.559
	θ_4	X_t	0.939	8.648
	θ_5	BKI_t	0.119	5.429
	α_0		14623.9	3.689
	α_1		0.816	2.762
(2)	θ_6	Const.	-537.601	-3.773
	θ_7	P_t	-697.158	-3.345
	θ_8	P^*_{t-1}	889.720	4.192
	θ_9	M_t	0.551	6.253
	θ_{10}	AKO_t	0.034	2.815
	α_0		3336.38	2.423
	α_1		0.695	2.115
	α_2		0.390	2.661
(3)	θ_{11}	Const.	486.84	2.147
	θ_{12}	P_t	282.00	1.867
	θ_{13}	P^*_{t+1}	-357.97	-1.888
	θ_{14}	r_t	-79.632	-2.422
	θ_{15}	NLH_t	0.396	4.704
	ρ_3		0.823	7.315

Table 1 (to be continued on next page)

Table 1 (Continued from previous page)

Equation	Coefficient	Variable	Estimate	Asymp. t Value
(4)	θ_{16}	Const.	505.062	11.108
	θ_{17}	P_t	-320.886	-1.742
	θ_{18}	P^*_{t+1}	219.211	1.804
	θ_{19}	r_t	-40.058	-1.982
	θ_{20}	NSH_t	0.572	19.436
	α_0		1336.29	1.058
	β_1		0.737	3.159
(5)	θ_{21}	Const.	-0.035	-3.783
	θ_{22}	P_t	1.032	299.745
	θ_{23}	X_t	0.342×10^{-4}	4.163
	θ_{24}	IA_t	-0.300×10^{-2}	-12.738
	α_0		0.264×10^{-4}	2.173
	α_1		0.789×10^{-2}	0.094
	α_2		0.983	2.411

TABLE 2

INTRA-SAMPLE SIMULATION

Variable	Correlation Coefficient	Theil's IC	% RMSE
P	0.9937	0.0129	2.7872
A	0.9975	0.0131	2.7905

TABLE 3

POST-SAMPLE SIMULATION

Variable	Correlation Coefficient	Theil's IC	% RMSE
P	0.9622	0.0097	1.9780
A	0.9903	0.0098	1.9649

TABLE 4

POST-SAMPLE FORECASTS OF THE SPOT RATE

Forecast	Correlation Coefficient	Theil's IC	% RMSE
AS	0.9903	0.0098	1.9649
P_{t-2}	0.7007	0.0294	5.8889
RWALK2	0.7105	0.0222	4.4471

ENDNOTES

1. Until 1988(5) the British Pound futures contract at the Chicago Mercantile Exchange International Monetary Market called for delivery of 25,000 Pounds and from 1988(6) onwards this contract called for delivery of 62,500 Pounds.

2. The British pound futures contract on the IMM calls for delivery in the months of March, June, September, December each year. Trading tends to concentrate in the near future, and hence futures rate quotations were selected according to the following rule:

   when the month is Jan., Feb. the future is Mar.;

   when the month is Mar., Apr., May the future is Jun.;

   when the month is Jun., Jul., Aug., the future is Sept.;

   when the month is Sept., Oct., Nov., the future is Dec.;

   when the month is Dec., the future is Mar.

3. This treatment, while consistent with the view that the risk premium is the difference between the expected spot rate and the current futures rate (e.g. Kaldor 1953, p. 23-28; Hodrick and Srivastava 1987, p. 19), is compatible also with the view that the risk premium should be treated as a separately observable variable, rather than as a residual (see Stein 1991, p. 39). In the absence of superior knowledge of such a variable, the risk premium has been treated here as the difference between private and public sector rates of return. It has, therefore, been assumed implicitly that the return required for uncertainty about private sector securities is the same as that required for uncertainty about unhedged foreign exchange positions.

4. Data on AKO are given in million US dollars in the Bank of England *Quarterly*

Bulletin, and have been converted to British pounds at the monthly average spot rate, taken from the *Federal Reserve Bulletin.*

5. Fuller (1976) has shown that the limit distribution of the t statistic for $\hat{\gamma}$ is independent of the number of lags of ΔZ in the equation.

6. Theil's inequality coefficient and per cent RMSE are defined in Pindyck and Rubinfeld (1981, pp.362, 364). Initial values for intra- and post-sample simulations are obtained by three stage least squares estimates of the structural parameters.

7. Futures prices employed in this model are, on average, two months from maturity, so that two month ahead forecasts by the structural model and the random walk model provide the most appropriate basis of comparison.

8. For both predictors, AS and P_{t-2} the unbiasedness hypothesis cannot be rejected. Random walk forecasts of the spot rate two months ahead were obtained by estimating the following model by OLS

$$A_t = \alpha + \beta A_{t-2}$$

where α , β are constants. From these estimates, fitted values $\hat{A}_t$ were obtained, which acted as forecasts.

9. This outcome, moreover, implies that agents do not form "fully rational expectations", and that Stein's (1986, pp. 71, 150-51) concept of "Asymptotically Rational Expectations" (ARE) would be relevant here. As emphasized in Section II, a test of the REH is a joint test of the expectations hypothesis, and the appropriateness of the economic model in question. In the case of Muth Rational Expectations (MRE), agents know the true economic model driving returns in practice. With ARE, the difference between the subjective distributions of agents and the objective distribution

of the system converges to zero with repeated sampling of information, so that ARE converge to MRE. The empirical results in this paper suggest that the model presented here combines and processes elements of information differently from the way that economic agents do, and as a consequence the model produces a superior forecast of the spot rate, compared with the forecast implicit in the lagged futures rate.

APPENDIX 1

UNIT ROOT TESTS

Variable	Model	Calculated ADF Statistic	5% Critical Value	Integration Order
SH	(8)	-4.1780	-3.4645	I(0)
LH	(8)	-7.1213	-3.4645	I(0)
SS	(8)	-6.7027	-3.4645	I(0)
LS	(8)	-5.2062	-3.4645	I(0)
A	(10)	-0.6386	-2.8967	I(1)
P	(10)	-0.6004	-2.8967	I(1)
NLH, NSH	(10)	-3.6724	-2.8967	I(0)
M	(9)	-3.0600	-3.4645	I(1)
X	(9)	-2.8520	-3.4645	I(1)
AKO	(9)	-5.3610	-3.4645	I(0)
BKI	(9)	-5.0545	-3.4645	I(0)
r	(9)	-4.5268	-3.4645	I(0)
IA	(9)	-3.1924	-3.4645	I(1)

APPENDIX 2

ENGLE-GRANGER COINTEGRATION TESTS

Equation	Variables	Calculated AEG Statistic	1% Critical Value
(1)	P_t, $\hat{P}_{t+1}$, X_t	-5.0265	-4.4732
(2)	P_t, $\hat{P}_{t+1}$, M_t	-5.1738	-4.4732
(3)	P_t, $\hat{P}_{t+1}$	-5.5255	-4.0330
(4)	P_t, $\hat{P}_{t+1}$	-5.5255	-4.0330
(5)	A_t, P_t, X_t, IA_t	-7.5588	-4.8677

Appendix 3

Johansen Cointegration Procedure: λ Max Test

Equation	Test Statistic	5% Critical Value	No. of Coint. Vectors: m
(1)	50.3715 21.1027	42.44 25.32	m = 0 m ≤ 1
(2)	49.1276 15.9388	42.44 25.32	m = 0 m ≤ 1
(5)	78.2362 40.1550	62.99 42.44	m = 0 m ≤ 1

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