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CONTENTS

Abstract

The overseas sector holds financial assets denominated in a wide range of currencies in various financial centres around the world. In this paper we analyse changes in the composition of sterling and foreign currency denominated assets held in the UK by the overseas sector. Although these assets constitute a relatively small proportion of the total financial wealth of the overseas sector nevertheless movements between sterling and foreign currency assets held in the UK may be reflected in other financial centres. However, in this paper (because of data limitations) we only attempt to model asset switching within the UK market and institutions.

In the model we assume that overseas residents (or their portfolio managers in London) first decide on the split between total asset holdings denominated in sterling and in foreign currencies: in this study the latter consist mainly of dollar deposits held in UK banks. In the 'second stage' of their decision process overseas residents decide on the allocation of their total sterling asset holdings between (net) sterling bank deposits, UK company securities and UK public sector debt.

The model which underlies these choices is based on consumer theory and the restrictions implied by the theory (eg homogeneity, symmetry) are tested and found to hold. Adjustment of actual asset holdings to their (long-run) desired levels is not instantaneous and depends on disequilibria in other asset holdings.

We find that there is substantial switching between total sterling assets and foreign asset (held in the UK) in response to a change in expected rates of return. However, such changes in asset holdings are in response to a long-term view of relative rates of return, suggesting hedging behaviour and risk aversion by some agents. Thus although speculation probably takes place in the very short-term, fund managers are likely to 'close-out' their speculative positions over a time period shorter than one quarter, the frequency of data used in the study.

The choice within the set of sterling assets is influenced by relative rates of return. UK company securities and (net) sterling deposits (held in UK banks) are substitutes as are public sector assets and UK company securities.

Overall the results indicate that broad (quarterly) movements of assets held in the UK, by overseas residents may be explained by a model in which overseas residents respond in a predictable manner to changes in expected rates of return.

1 INTRODUCTION

Without intervention and in the absence of capital controls the (bilateral) exchange rate may be viewed as the 'price' which ensures a zero net asset demand for a currency. Bilateral exchange rates are determined in an interdependent system which is often presented in terms of a portfolio balance model (Brainard and Tobin 1968, Branson et al 1977, 1979). The flex-price monetary model, FLPM (Frenkel 1976, Bilson 1978) and the sticky-price monetary model, SPM (Dornbusch 1976, Buiter and Miller 1981) may be shown to be special cases of the portfolio balance approach. The efficient markets hypothesis (EMH) is often combined with the assumption of perfect capital mobility and risk neutrality, when discussing the behaviour of the foreign exchange market (eg MacDonald 1988, Boothe and Longworth 1986). Recently this approach has been extended to include the modelling of time varying risk premia (Taylor 1988, Giovanni and Jorion 1987, 1988). The incorporation of 'news' in models based on the EMH have also been used to explain the volatility of exchange rates (MacDonald 1988).

An understanding of exchange rate determination is clearly of policy importance. Governments and Central Banks alter domestic interest rates in an attempt to influence exchange rates or intervene directly in the foreign exchange market. In the UK, with the ending of sole reliance on quantitative targets for the money supply, the exchange rate has become an important indicator of monetary conditions (Goodhart 1989). Within the portfolio balance framework, Central Banks may influence the exchange rate either by altering asset supplies, (eg by open market operations) or by a change in the budget deficit or by directly influencing expected returns or perceptions of risk.

In part, because of the failure of existing models of exchange rate determination, an interesting strand of the theoretical literature deals with the implications of rational bubbles (Blanchard 1979, Frankel and Froot 1986) and noise traders (Black 1986, Campbell and Kyle 1986, De Long et al 1988). In both of these approaches price movements may not be determined entirely by 'economic fundamentals'. Empirical work on these two strands of the literature is in its infancy (eg Goodhart 1988, Garnett and Taylor 1989).

In the empirical literature there are two broad approaches to analysing the exchange rate: the 'reduced form' and the 'structural approach'. The structural approach we define as estimation of asset demand and supply functions by different sectors in the economy and then solving the resulting system for the equilibrium exchange rate (simultaneously with other asset prices). Estimation of 'complete' models of this type are rare, in part because of acute data requirements but Green (1984), Keating (1985) and Kearney and MacDonald (1985, 1986) provide recent examples. The literature on the various reduced form approaches is voluminous (see MacDonald (1988) for survey) but what distinguishes them from the structural approach is that the exchange rate (either real or nominal) is usually taken to be the dependent variable.

Empirical results from these two broad approaches can only be described as 'mixed'. There is certainly no 'consensus' model of exchange rate determination and it is clear that research will continue on a number of fronts simultaneously. We do not wish to take an entrenched view on what is the best approach given theory and empirical evidence, todate: considerable uncertainty surrounds any modelling strategy in this area. In this paper we add to the portfolio balance literature by examining the demand for assets held in the UK by the 'Overseas Sector'. The demand functions for this sector are required if one is to 'close' a model of the UK financial sector where asset demands of the domestic sectors have already been modelled. The overseas sector's asset demands are often neglected but are clearly of importance in determining capital flows and influencing exchange rate movements. Our model is closest in structure to that of Kearney and MacDonald (1985) who also utilise a portfolio balance model albeit for the UK non-bank private sector NBPS. They use a four asset model (namely money, domestic bonds, foreign assets and bank lending). They define wealth in terms of the above assets and hence impose adding up constraints on the demand functions which also have (untested) unit wealth elasticities imposed and dynamics are incorporated via interdependent asset adjustment. They find that unrestricted estimates yield 'wrong signs' on 11 (out of 16) interest rate coefficients. Use of the Theil-Goldberger 'pure and mixed' estimator reduces the number of wrong signs to 5 (out of 16). Green (1984) finds 57 per cent of 'wrong signed' own rates in the demand functions for overseas assets while Keating (1985) utilises a highly constrained mean-variance model (eg a constant, scalar diagonal covariance matrix) and imposes some coefficients. A drawback of the mean-variance model is that it does not allow considerations of hedging and asset demands are determined by one period ahead expected returns (and a constant coefficient of risk aversion and variance-covariance matrix of asset returns). Lewis (1988) estimates a portfolio balance model using data disaggregated by currency (for aggregate bond holdings) but finds that the substitution parameters are 'imprecise'. The above empirical studies have undoubtedly contributed to our understanding of the demand for foreign assets but we hope to improve on these approaches. In this paper we use a systems approach where the theoretical structure is based on the Almost Ideal Demand System (AIDS) (Deaton and Muellbauer 1980). This enables us to remain agnostic on the issue of hedging demand. Unlike Green (1984), Kearney and MacDonald (1985) and Keating (1985) we test the theoretical restrictions of homogeneity, symmetry and homotheticity. Dynamics are modelled in an interdependent error feedback system, which provides a useful interpretation in terms of asset disequilibrium. Our approach has its limitations and like other portfolio studies we do not incorporate explicit time varying measures of risk. However the empirical results are encouraging and are an improvement on previous systems approaches and on results from 'single equation' currency substitution models (eg Spinelli (1983)).

The rest of this paper is organised as follows. In Section II we outline the theoretical model and in Section III we consider the modelling of short-run dynamics in a systems framework and associated econometric problems. In Section IV we discuss data problems and in Section V we present our empirical results. We conclude with a brief summary.

II THE AIDS MODEL

The representative agent is assumed to distribute his wealth between alternative assets in order to minimise the cost of achieving a given level of utility.[1] The axioms of rational choice in demand theory (ie the existence of consistent preferences) are met providing we choose a cost function that is concave and homogeneous of degree one in prices.[2] Of the several flexible functional forms available we use the PIGLOG (Price Independent Generalised Logarithmic) which, in common with others (eg indirect translog, Christensen et al 1975) is a second order approximation but has the advantage of desirable aggregation properties. Within the PIGLOG class we use the AIDS cost function (Deaton and Muellbauer 1980).

The budget constraint is:

$$\sum_i p_{it}^{\tau} \, a_{it+1}^{\tau} = W_t^{\ \tau} \tag{1}$$

where

p_{it}^{τ} = 'real price' $= [\,(1 + r_{it}\,)\ (1 - g_z)\,]^{-1}$

r_{it} = <u>expected</u> (proportionate) nominal return on asset i, between t and t+1 (including any capital gains).

g_z = <u>expected</u> (proportionate) rate of goods price inflation between t and t+1.

(1) One can argue that ultimately utility depends only on current and future consumption. However, in the absence of fully contingent binding contracts, when saving takes place, agents must hold some asset stocks, and it is reasonable to assume that agents are not indifferent to the composition of their assets. Barnett (1980) incorporates assets and goods in the utility function in deriving Divisia aggregates. Asset holdings represent purchasing power over future consumption goods. In this framework any factors not in the wealth constraint that influence the composition of asset holdings, will implicitly be captured in the parameters of the utility function. (See also Taylor and Clements 1983.) An alternative approach as to why agents hold both cash and non-interest bearing chequing accounts may be found in 'cash in advance' models (Lucas 1984, Hartley 1988).

(2) The cost function is also usually assumed to be continuous in prices and that the first and second derivates with respect to prices exist.

lnp_{it} = nominal price (= $ln(1 + r_{it})^{-1} \approx - r_{it}$)

a^{τ}_{it+1} = real asset holdings (= a_{it+1}/z_{t+1})

a_{it} = nominal asset holdings of the i^{th} asset

z_t = goods price index

W_t^{τ} = real wealth (= W_t / Z_t)

P^{τ}_{it} is a real discount factor but by analogy with the AIDS model applied to goods, we designate it as a real AIDS price.

Solving the constrained cost minimisation problem leads to asset shares, equations (see for example, Barr and Cuthbertson 1989, Weale 1986) depending on the AIDS prices, $ln\, p^{t}_{jt}$ and real wealth:

$$s_i = \alpha_i + \sum_j \gamma_{ij}\, lnp^{\tau}_{jt} + \beta_i\, ln(W^{\tau}/P^{*\tau})_t \tag{2}$$

where

$$s_i = a_{it} / W_t \tag{2a}$$

$$ln\, P_t^{*\tau} = \sum s_i\, lnp^{\tau}_{it} \tag{2b}$$

Note that $ln\, P_t^{*\tau}$ may be interpreted as a composite real discount factor.

The long-run AIDS share equations in vector notation are:

$$s^{*}_t = \Pi X_t = \Gamma\, lnp^{\tau}_t + B\ RW_t + \alpha \tag{3}$$

s^{*}_t = k x 1 vector of asset shares

lnp^{τ}_t = k x 1 vector of real asset prices

RW_t = log of real (AIDS) wealth(= $ln\,(W^{\tau}/P^{*\tau})_t$)

Γ = (k x k) matrix of 'price' coefficients (γ_{ij})

B = (k x 1) vector of real wealth coefficients (β_i)

α = k x 1 vector of constants

Π = $[\ \Gamma : B : \alpha :]$

X_t = $[\ ln\, p^{\tau}_t, RW_t, 1\ ]$

The theoretical restrictions of the AIDS model based on consumer theory are:

The adding up constraints

$$\sum_i \alpha_i = 1, \ \sum_i \gamma_{ij} = 0, \ \sum_i \beta_i = 0 \tag{3a}$$

Homogeneity:

$$\sum_j \gamma_{ij} = 0 \tag{3b}$$

If homogeneity holds then the share equations (2) may be written in terms of nominal <u>relative</u> prices, and inflation if included separately should be statistically insignificant.

Symmetry and negativity (of the Hicksian demand functions) are direct consequences of the axioms of rational choice. The former implies:[1]

$$\gamma_{ij} = \gamma_{ji} \tag{3c}$$

Negativity arises from the concavity of the cost function and implies that the matrix of coefficients k_{ij}:

$$k_{ij} = \gamma_{ij} + \beta_i \ \beta_j \, ln(W^{\tau}/p^{\tau}) - s_i \, \delta_{ij} + s_i \, s_j \tag{3d}$$

is negative semi-definite (δ_{ij} is the Kronecker delta).

Thus our systems approach implicitly imposes data admissability in the form of adding up constraints and the additional theoretical constraints of symmetry homogeneity and negativity. If symmetry and homogeneity hold then this reduces the number of parameters to be estimated and increases efficiency. In addition one might wish to judge the model on more intuitive notions, (eg that own price effects are negative, wealth and expenditure elasticities are 'reasonable' etc).

The wealth and compensated own-price and cross-price elasticities are:

$$E_w = (\beta_i / s_i) + 1 \tag{4a}$$

$$E_{ij}(p) = (s_i)^{-1} \, k_{ij} \tag{4b}$$

Semi-elasticities with respect to the <u>annualised percentage</u> nominal yield R (ie R = 400r) are given by

$$E_{ij}(R) = E_{ij}(p)/4 \tag{4c}$$

III SEPARABILITY AND TWO-STAGE BUDGETING

It is clearly somewhat impractical to simultaneously estimate all asset demands of the overseas sector. To delineate the problem we only consider holdings of highly 'liquid assets', namely net foreign currency deposits (NFD), net

(1) Note that 'adding up' and symmetry imply homogeneity. (Although homogeneity and 'adding up' do not imply symmetry.)

sterling deposits (NSD), UK company securities (CS) and UK public sector long-term debt (PSL). Apart from direct investment, these categories constitute the major asset holdings of the overseas sector in the UK. In more formal terms we assume weak separability (Deaton and Muellbauer 1980) between our chosen four assets and other assets held by the overseas sector. We realise this is a strong assumption but if it is widely at variance with the data this will be reflected in 'poor' parameter estimates. Total wealth held in the four assets is therefore assumed to be predetermined. We further delineate the problem by assuming weak separability between holdings of foreign assets (NFD), and the set of three sterling assets (NSD, CS, PSL): this constitutes the 'upper level' decision in the two-stage budgeting process. Having decided on the upper level decision the overseas sector is then assumed to distribute its wealth held in sterling assets between NSD, CS and PSL and this is the 'lower level' or second stage decision.

A word of caution is in order here concerning the two-stage budgeting procedure proposed. Weak separability is necessary and sufficient for the <u>second</u> stage (ie 'lower level' decisions) of two-stage budgeting. However weak separability has some drawbacks. First, it places quite severe restrictions on the degree of substitutability between goods in different groups (Pudney 1981). For example whole groups will be substitutes or complements with each other. The second problem is, potentially, more serious however. In modelling the 'upper-level' decisions it would be extremely useful to be able to establish the maximisation problem in terms of <u>group</u> price and quantity indices (rather than having to utilise <u>all</u> prices that the agent faces), since this would considerably reduce the number of parameters to be estimated for the 'upper-level' equations. Strictly speaking this is only possible under somewhat restrictive assumptions. If preferences are homothetic which implies all 'expenditure elasticities' are unity or equivalently that the budget shares <u>within</u> each group are independent of total group expenditure, then 'group price indices' can be legitimately used in determining the upper level allocations. If preferences are <u>not</u> homothetic then group price indices can be used providing the utility function is strongly separable and has the Generalised Gorman polar form. However the former implies 'additivity' between groups:

$$u = u_a(q_a) + u_b(q_b) + \ldots$$

Additivity is restrictive in that (a) inferior goods are ruled out; (b) goods can only be substitutes given that inferior goods are not allowed; (c) expenditure elasticities are proportional to price elasticities. If neither of the above assumptions holds then the group price indices P_a say are given by $P_a = P_a(u_a, p^a)$ and are dependent on u_a, which in turn depends on <u>all</u> other prices outside the group. One possibility is not to invoke homotheticity and assume that P_a does not vary very much with u_a and hence most of the 'explanation' of P_a is the sub-set of prices, p^a. Given the other approximations involved in empirical studies this may be a reasonable expedient to adopt if homotheticity is not found to hold in the data. However, in our study homotheticity holds in the long run and hence we can legitimately apply the two-stage budgeting approach.

We also implicitly assume weak intertemporal separability.(1) Formal non-parametric tests of separability (Varian 1983, Swofford et al 1986) are beyond the scope of the present paper but we do provide an ad-hoc test for separability.

IV DYNAMIC ADJUSTMENT AND ECONOMETRIC ISSUES

In a system of asset demand equations if we include only own-lags, then we must implicitly accept that all assets adjust at the same rate (Smith 1975). To avoid this problem cross-lagged terms must also be included (Brainard and Tobin 1968). The latter can be rationalised by generalising the quadratic cost of adjustment function of Christofides (1976):

$$L^* = (s - s^*)_t' \, C_1 \, (s - s^*)_t + \Delta s_t' \, (C_2) \, \Delta s_t - (\Delta s_t) \, C_3 \, (\Delta s_t^*) \qquad (5)$$

(1) For AIDS models that relax this assumption see Weissenberger (1986) and Rossi (1987).

where C_i (i = 1, 2, 3) are conformable adjustment matrices. Minimising L^* with respect to s_t subject to the budget constraint we obtain generalised error feedback equations:

$$\Delta s_t = \Pi^* \, \Delta X_t + L\,(s - s^*)_{t-1} \tag{6}$$

where the disequilibria in (k-1) asset shares at time t-1 influence the current period adjustment of any particular asset share. Since $\sum_{1}^{k} (s_i - s_i^{\,*})_{t-1} = 0$, only the (k - 1) independent disequilibrium shares are required in (Anderson and Blundell 1983) (6) (ie L is (k x (k-1)). The adding up restrictions imply that the columns of Π^* and L sum to zero. In addition on intuitive grounds we might expect the diagonal elements of L to be negative. However the latter is not required for dynamic stability. As long as the eigenvalues of the appropriate adjustment matrix have modulus less than unity, then the system is dynamically stable.(1)

The above adjustment equation could be further generalised by including additional lagged own and cross-shares that is, terms in Δs_{t-j} and ΔX_{t-j}; although the dynamics are then not consistent with the quadratic cost of adjustment function (5). Indeed one can also interpret (6) merely as a reasonably parsimonious method of incorporating dynamics while maintaining the adding up restrictions but without alluding to the cost function (5). We feel that quadratic costs are very unlikely to apply to adjustments in financial assets (eg costs are unlikely to rise smoothly with the size of the transaction undertaken) and we therefore interpret equation (6) merely as a convenient method of characterising sluggish adjustment (Hendry et al 1984). In principle it is possible to generalise (5) to yield a <u>multiperiod</u> interdependent costs of adjustment function but this rapidly becomes intractable (Currie and Kennally 1985), and still retains the unrealistic quadratic form.

Drawing on (3) and (6), the short-run equations may be represented by the following interdependent EFE:

$$\Delta s_t = C\Delta \, lnp_t^{\tau} + K\Delta RW_t + L(s - s^*)_{t-1} \tag{7}$$

where C, K and L are suitably dimensioned matrices of short-run parameters. Substituting for s_{t-1}^* from (3) in (7) then gives our set of non-linear estimation equations. The advantage of (7) over an unrestricted ADL equation is that we can easily test and impose restrictions on the long-run parameters (eg homogeneity, symmetry).(2) In contrast, Kearney and MacDonald (1985), Green (1984) and Weale (1986) do not test for <u>long-run</u> symmetry and homgeneity and these properties do not hold in the long run, about which economic theory is thought to be more informative. We are also able to test the long-run parameters obtained to see if they form a co-integrating vector (Hall 1986, Hendry 1986, Granger 1986).

We use a general to specific modelling strategy with equation (7) as our initial maintained hypothesis without any restrictions on the parameter matrices C, K, Γ, B (other than the adding-up restrictions to obtain data admissability). The theoretical structure imposed by adopting a systems approach therefore limits the extent to which one can indulge in 'overfitting' and data-mining. The final parsimonious system of equations is subject to the usual test procedures (although understandably these are not as numerous as found in single equation studies).

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(1) If the disequilibrium term for asset 1 is excluded then the <u>estimated</u> adjustment matrix is:
$L = (\underline{1}_2, \underline{1}_3 \ldots \underline{1}_k)$ $\quad k \times (k-1)$
The dynamics of the full model may be written
$s_t = (I_k + L^*)\, s_{t-1}$
Where
s_t is $k \times 1$
$L^* = (i, \underline{1}_2, \underline{1}_3 \ldots \underline{1}_k)$ is $(k \times k)$
$i = (1, 0 \ldots 0)$ is $(k \times 1)$
$I_k = k \times k$ identity matrix
One of the eigenvalues of $(I_k + L^*)$ is unity and stability requires that the other (k - 1) eigenvalues have negative real parts.

(2) However one of the practical disadvantages for our particular model is the need to provide 'good' starting values for the parameters, to ensure convergence in the non-linear system. An alternative is to use the procedure in Bewley (1979).

The equations are estimated by 3SLS (Zellner and Theil 1962) treating current period prices and wealth as endogenous. The instruments used are two lagged values (t-1, t-2) of all prices, wealth and asset shares. Corrections for serial correlation in systems of equations are not possible with our current software (Berndt and Savin 1975) but because of our flexible lag response this was not found to be an acute practical problem.

In heavily parameterised systems estimated with a relatively modest number of observations, there is likely to be over-rejection of null hypotheses if asymptotically valid test statistics are used (eg Deaton 1974, Bewley 1983). In general it appears that over-rejection becomes more severe as the number of equations in the system increases or as the sample size becomes smaller. Small sample adjustments are problematic in dynamic simultaneous systems but following Mishkin (1983) we use:

$$WA(r) = [\,(nT - k)/nT\,]W \tag{8}$$

where:

W = Wald statistic

n = number of independent equations

T = total number of observations

r = number of parameter restrictions

k = number of estimated parameters in the system

$WA(r)$ is approximately distributed as $\chi^2(r)$ under the null.(1) Pudney (1981) also suggests upward adjustment of critical values (hence making rejection of the null more likely) but we have not pursued this.

In testing parameter constancy we have used a systems analogue to the Salkever (1976) test which allows the use of a fixed instrument set and fixed variance covariance matrix. For testing parameter constancy over r additional periods we augment each equation with r (. . .0, 1, 0 . . .) dummies. The coefficients and t-statistics on the dummies then yield estimates of the outside sample forecast errors and their statistical significance. A Wald (or adjusted Wald) test on all the dummies yields an asymptotically valid test of the parameter stability in the system as a whole. The dummies are included in the instrument set.

V DATA USED AND ASSET CLASSIFICATIONS

The asset categories modelled of the overseas sector are:

NSD =net sterling deposits (ie sterling sight plus time deposits less sterling loans)

CS =UK company securities

PSL = public sector long-term debt

NFD = net foreign currency deposits (ie foreign currency deposits less foreign currency loans)

(1) In our model for the 'lower level' decision the number of regressors is relatively large and the adjustment factor is of the order of 0.6

Since there is considerable matching of both foreign currency and sterling deposits and loans, the net position in these assets seemed the appropriate decision variable (Lewis and Davis 1987).

The flow data is taken from the 'Flow of Funds' matrix in Financial Statistics. Revaluation indices are chosen to be consistent across sectors of the complete matrix. Benchmark stocks are then chosen such that <u>all</u> elements in the matrix satisfy the accounting identities (ie zero row-sums and column sums equal to the NAFA). The data on company sector assets used here therefore comes from a fully consistent complete stock-flow matrix.

The rate of return on *NSD* is the rate on 'parallel money-market' assets, namely the 3-month rate on Local Authority bills. Various rates of return for *CS*, *PSL* and *NFD* were tried all of which included the running yield and capital gains. One quarter and one year ahead capital gains did not perform well and eventually we used the running yield plus a 3-year backward looking capital gain. The F-T Actuaries price index for all government stock and the running yield are used for *PSL*. The dividend yield and F-T Actuaries all share index are used for *CS*, while the return on *FCD* is the yield on 3-month dollar deposits in London plus capital gains due to changes in the dollar-sterling exchange rate. We also used a weighted exchange rate consisting of the Dollar, Yen and DM and results were marginally inferior. But as over 75 per cent of foreign currency deposits and loans are held in dollars in London, the use of the dollar-sterling rate seems reasonable (Lewis and Davis 1987).

All data used are seasonally unadjusted but seasonal dummy coefficients are not reported. The regressions are run over the period after the ending of exchange controls and critical values of test statistics are given at a 5 per cent significance level (unless stated otherwise).

In making its 'upper level' decision the overseas sector allocates it wealth between total sterling assets *SA* (ie *NSD+CS+PSL*) and total foreign assets,[1] *NFD*. The share of sterling assets s_1^u is:

$$s_1^u = (SA/W) \tag{9a}$$

$$W = SA + NFD \tag{9b}$$

Under symmetry (or homogeneity) the upper level dynamic AIDS share equation for sterling assets, s_1^u is:

$$\Delta s_{1t} = c_{11}\,\Delta ln\,(p_1^u/p_2^u)_t + k_1\,\Delta ln\,(W^*/P^*)_t + a_{11}\,ln\,(p_1^u/p_2^u)_{t-1} + b_1\,ln(W/P^*)_{t-1} + L_{11}\,s_{1t-1}^u \tag{10}$$

where:

W^* = $(W/Z^*)$where Z^* =world goods price index (ie weighted average of OECD consumer price indices)

lnp_1^u = weighted average of the nominal rates of return on the three sterling assets *NSD*, *CS* and *PSL* (see section III) (ie an AIDS composite price index using mean shares as weights)

lnp_2^u = nominal return on 3-month dollar deposits in London adjusted for exchange rate changes (\$/£), expressed as an AIDS price

(1) The wealth variable for the upper level decision is wealth of the overseas sector held in the UK. It could be argued that a wider measure of wealth is required but such data is not readily available. However, our real wealth variable is highly correlated with world GNP (simple correlation coefficient = 0.93), a scale variable frequently used in single equation demand functions for 'foreign assets'. Also it is probably the case that fund managers deal predominantly with a particular financial centre because of established broker-customer relationships and hence will switch between assets denominated in different currencies within one financial centre. This is tantamount to assuming a preferred 'geographical habitat' and strengthens the case for our particular scale variable. Also 'London is probably more of an international centre than either New York or Tokyo' (Lewis and Davis 1987, p 234) and hence this should help minimise any measurement error in our chosen wealth variable. Masson (1987), provides an informative discussion of practical problems in choosing a suitable wealth variable in a portfolio balance model.

lnP^{-s} $= (s_1^{\text{u}}\ ln\ p_1^{\text{u}} + s_2^{\text{u}}\ ln\ p_2^{\text{u}} - g_z)$, aggregate AIDS real price. Where g_z = rate of world price inflation. Sample means are used for s_i^{u}.

The demand for foreign assets is determined by the budget constraint $s_2^{\text{u}} = (1 - s_1^{\text{u}})$.

The share equations for the three 'lower level' sterling assets are

$$s_t = C\ \Delta lnp_t + K\ \Delta ln(W^s/P^s)_t + L\ [s_{t-1} - \Gamma\ lnp_{t-1} - B\ ln(W^s/P^s)_{t-1}] \quad (11)$$

where

$s_{1t} = (NSD/W^s)_t$

$s_{2t} = (CS/W^s)_t$

$s_{3t} = (PSL/W^s)_t$

$W_t^s = [(NSD + CS + PSL)/Z]_t$

$s_t = (s_{2t}, s_{3t})'$

$lnp_t = (lnp_{1t}, lnp_{2t}, lnp_{3t})'$

$lnP_t^s = \sum s\ lnp_{it} - g_z^s$

Z_t =sterling goods price (TFE deflator)

g_z^s =rate of inflation of sterling goods prices

Equation (11) is estimated as a (2x2) system with the share of 'asset 1' being determined by the budget constraint. Since total sterling assets are assumed to be weakly separable from foreign assets, the appropriate inflation variable uses an aggregate sterling goods price index. The terms in square brackets in (11) represent the long-run disequilibria in asset shares $(s - s^*)_{t-1}$, with adjustment matrix 'L'.

VI EMPIRICAL RESULTS

A Upper Level Decision: Total Sterling and Foreign Currency Assets

The preferred 'upper level' share equation for total sterling assets with short-run and long-run symmetry imposed is:

Total sterling assets

$$\Delta s_{1t}^{\text{u}} = \underset{(1.6)}{0.11} - \underset{(1.8)}{1.35}\ \Delta ln(p_1^{\text{u}}/p_2^{\text{u}})_t - \underset{(2.6)}{0.99}\ ln\ (p_1^{\text{u}}/p_2^{\text{u}})_{t-1} - \underset{(5.9)}{0.22}\ s_{1t-1}^{\text{u}} \quad (12)$$

IV, 80(1)-86(4), $R^2 = 0.69$, DW = 1.9, BP(4) = 8.2, ARCH(1) = 2.2, RAM(1) = 0.8

The long-run semi-elasticity of sterling asset holdings with respect to a one per cent change in the annual yield (on sterling or foreign assets) is -1.63. However, because the rate of return includes capital gains the change in asset holdings can be substantial: a change in the relative price variable equal to one standard deviation would lead to a

12 per cent change in sterling asset holdings. Thus there is substantial switching between sterling and foreign currency assets in response to changes in relative rates of return. However, the response is far from infinite as suggested by the empirical work on UIP. It may be that market segmentation is at work: hedging behaviour by some asset holders takes place based on a long-run view of expected yields while short-term speculation ensures UIP.[(1)] In quarterly data we are only able to pick up the behaviour of the former.

The Box-Pierce statistics indicate the absence of serial correlation of up to order 4 (critical value for BP(4) is 9.5). The ARCH (Engle 1982) and Ramsey (1969) RESET tests (with critical values 3.8 and 1.7 respectively) indicate the absence of autoregressive condition heteroscedasticity and mis-specified functional form.

The parameter stability of equation (12) is shown in Table 1. The Chow test indicates parameter stability over a number of sub-periods and the individual coefficient values can be seen to be reasonably constant.

The above equations have symmetry (and homogeneity) and homotheticity imposed. Adding a lagged level of real wealth to the above equation over the above sample periods (in Table 1) never yields a *t*-statistic above 1.5 (and all but one of the *t*-statistics are less than unity). Homotheticity is therefore not rejected. We test for long-run and short-run symmetry and homogeneity in (12) by including the level and first difference of both a nominal price variable and the rate of inflation. [(2)] The quasi-likelihood ratio test yields QLR(4) = 2.6(χ^2_c = 9.5) and we therefore cannot reject these restrictions.

B Lower level decision: sterling assets: NSD, CS, PSL

In an EFE (equation 7) without homogeneity and symmetry imposed, long-run and short-run own-price effects are all negative. We can easily accept the hypothesis of long-run, homogeneity W(2) = 0.6 (χ^2_c = 6.0) and the hypothesis of long-run symmetry (which implies homogeneity) W(3) = 1.2 (χ^2_c = 7.8). The model therefore satisfies two of the basic axioms of rational choice in demand theory. Short-run homogeneity W(2) = 4.0 (conditional on long-run symmetry and homogeneity holding) and short-run homogeneity plus symmetry W(3) = 4.2 are also not rejected on a Wald test.

In the next stage of our simplification procedure we therefore impose long-run and short-run symmetry and homogeneity. We then sequentially exclude coefficients that have both low t-statistics and a relatively small numerical impact (Table 2). We can accept long-run homotheticity (B2=B3=0) and the zero restrictions K3=L23=0 on a Wald test, W(4)=7.7 (χ^2_c = 9.5).

The estimated parameters of the AIDS model (Table 2) are not very informative in terms of the usual economic concepts (eg elasticities) but the following general features are of interest. The 'own rate' long-run price coefficients, γ_{ii}, are negative and the γ_{ij} ($i \neq j$) for *CS* and *PSL* are positive indicating these assets are substitutes. There is weak complementarity between *NSD* and *PSL* although the long-run effect is not statistically significant. Finally *NSD* and *CS* are substitutes which does not seem unreasonable given that both are denominated in sterling but one is capital certain and the other capital uncertain.

The long-run semi-elasticities with respect to the annual rates of return (Table 3A) are shown together with the long-run impact on the absolute level of asset holdings (Table 3B). The own rate semi-elasticities of *NSD*, *CS* and *PSL* are 2.4, 4.9 and 2.7 respectively. It can also be seen from Table 3B (by looking down the columns R_1, R_2, R_3) that there is substantial switching between *NSD* and *CS* (Columns 1 and 2, Table 3B) and between *PSL* and *CS*

(1) Goodhart (1988) provides some evidence that hedging considerations are importan tin all but very short-term decisions involving currency exposure.

(2) Results are largely invariant to the use of world inflation because the variation in $\ln P^{w}$ and the individual real AIDS prices are dominated by the nominal price element ($\ln p_1^n$ $\ln p_2^n$). We prefe[r] a world inflation variable on *a priori* grounds when modelling the behaviour of the overseas sector because of its wider coverage. If Purchasing Power Parity (PPP) holds in the long-run result[s] will be invariant to the choice of the goods price numaraire.

(Columns 2 and 3, Table 3B). Asset switching is quite substantial relative to the quarterly mean flow into this set of assets (final column, Table 3B). In the long-run the effect of a £100 million increase in wealth held in sterling assets results in a fairly even distribution across the three assets (Column 4, Table 3B).

All the short-run price effects (Cij's) are smaller than their long-run values (Table 2B) and are well determined statistically.

The matrix of adjustment parameters (Table 2C) has all diagonal elements negative, indicating that excess holdings of asset 'i' leads to a fall in the share of asset 'i' in the subsequent period. The eigenvalues of the (augmented) L matrix have 'real parts' 0.87, 0.54 with zero imaginary parts, indicating a convergent response after a one quarter lag.

Further tests

Using the equations of Table 2 we perform tests for a first difference model, for weak separability, parameter stability and for a long-run co-integrating vector. We report each of these in turn and the results are summarised in Table 4.

The first difference model is a test that the remaining L_{ij} coefficients of Table 2C are jointly zero and is decisively rejected, W(3)=18.7, WA(3)=11.2 ($\chi_c^2 = 7.8$). The static model suffers from first order serial correlation which is indicative of mis-specified dynamics and can be rejected on this count alone.[1]

We performed a (somewhat ad-hoc) test for separability by including a lagged levels and first difference term in the 'price' of foreign assets lnp_2^u (ie the rate on 3-month dollar deposits adjusted for exchange rate changes) in our equations for sterling assets. These variables yield WA(4)=8.6 ($\chi_c = 9.5$) and therefore we can accept the null of separability. (There is only one coefficient with a t-statistic in excess of 1.7.)

Parameter stability is accepted over the period 1986(1)-1986(4) (WA(12)=15.0, $\chi_c^2 = 15.5$) but is rejected for the period 1979(1)-1979(4) [WA(18) = 21.8, $\chi_c^2 = 15.5$]. Although the long-run and short-run parameters are qualitatively unchanged when the data period is extended back to 1979(1), this structural break immediately after the ending of exchange controls may indicate a 'learning period' by agents.

We checked for stationarity of the residuals in our long-run equations. The Dickey Fuller (DF) tests are given in Table 4. If $\hat{\Pi}$ are the long-run parameter estimates, X the matrix of long-run variables and s the vector of asset shares,[2] the co-integrating residuals are $e = s - \hat{\Pi}X$. The test for stationarity in the residuals of equation i (i=1, 2, 3) uses:

$$\Delta e_{it} = \Theta_i e_{it-1} \tag{13}$$

The residuals are stationary for $\Theta_i < 0$ and $| t_\Theta | > 3.6$ (at a 10 per cent significance level), but critical values are only approximate (Engle and Yoo 1987). For our equations the test statistics (Table 4) are somewhat below their critical values although the point-estimates of Θ_i are always negative.[3] But given that the tests have low power against highly dynamic stationary alternatives (Granger 1986, Engle and Granger 1987) and given our small sample, the

(1) The Box-Pierce, BP(1), statistics for the residuals from the static regressions for *NSD*, *CS* and *PSL* are 6.2, 8.2 and 2.9.

(2) Asset shares are found to be I(1) in the data set as are the AIDS price variables. As shares are bounded (0,1), they cannot be I(1) with a Guassian error near the boundary point. Similar considerations apply to a variable such as the per cent unemployed or a bilateral exchange rate (which is bounded below).

(3) Higher order lags Δe_{it-j} (j = 1, 2, . 4) are not significant in equation 13.

results can at best be taken to be indicative of stationary residuals and the existence of a co-integrating set of parameters (which may not be unique, Johansen 1988).(1)

The statistical significance of the individual parameter estimates particularly those on the price variables are encouraging. The R^2 for the preferred equations are reasonable given we are explaining asset shares (Table 2C). The Box-Pierce statistics indicate the absence of serial correlation of up to order 4. Thus on statistical and economic grounds the 'lower level' equations for the choice between the three sterling assets are acceptable.

VII SUMMARY

We have invoked separability and two-stage budgeting to model the demand for foreign and domestic assets held in the UK by the overseas sector within a coherent theoretical framework provided by the AIDS model. Dynamics are adequately modelled using a systems error feedback approach which allows tests of and restrictions to be imposed on long-run parameters across different equations. The estimated model satisfies the theoretical restrictions of homogeneity, symmetry and homotheticity, broadly characterises the data and would appear to yield results which are an improvement on previous work in this area.

l price element (lnp_1^* lnp_2^*). We prefer
arity (PPP) holds in the long-run results

(1) The distribution of test statistics in multiple equations with I(1) variables with multiple unit roots has not yet been fully developed in the literature (see for example West 1988). This makes inference in time series models with potential I(1) variables somewhat hazardous.

Table 1: Upper Level: Share of Total Sterling Assets, Δs^u_{1t}

Estimation Period	Constant	$\Delta \ln (p_1^u/p_2^u)_t$	$\ln (p_1^u/p_2^u)_{t-1}$	s^u_{1t-1}
80(1)-84(4)	0.05 (0.7)	-0.88 (1.6)	-0.95 (2.6)	-0.11 (7.2)
80(1)-85(4)	0.12 (2.0)	-0.94 (1.6)	-1.15 (3.0)	-0.23 (6.8)
80(1)-86(4)	0.11 (1.6)	-1.35 (1.8)	-0.99 (2.6)	-0.22 (5.9)
79(1)-86(4)	0.10 (1.8)	-0.71 (1.1)	-0.71 (2.7)	-0.20 (7.6)
78(1)-86(4)	0.05 (1.1)	-0.61 (0.9)	-0.38 (1.3)	-0.11 (10.4)

Outside-Sample Stability Tests

	Chow test[1]	Critical Value
86(1)-86(4)	2.8	2.9
85(1)-85(4)	2.1	3.0
79(1)-79(4)	0.1	2.8
78(1)-78(4)	1.6	2.7

(1) The Chow test is calculated using the consistent estimates of the residuals from the IV regressions. As in the usual stochastic regressors case it is therefore only approximately distributed as a F-distribution.

Table 2: Preferred Error Feedback Equation, Lower Level (LR and SR Symmetry and Homogeneity Imposed)[1]

A Long-run parameters

	Price matrix Γ		
	$\ln p_1$	$\ln p_2$	$\ln p_3$
1. NSD	-2.91 (1.8)	3.76 (1.6)	-0.85 (0.8)
2. CS	3.76	-7.2 (1.9)	3.44 (1.8)
3. PSL	-0.85	3.44	-2.59 (2.2)

B Short-run parameters

	Price matrix C			Wealth
	Δg_t	$\Delta \ln p_2$	$\Delta \ln p_3$	ΔRW
1. NSD	-1.98 (3.7)	1.01(2.8)	0.97(2.7)	+0.26 (4.0)
2. CS	1.01	-1.58(5.0)	0.57(2.2)	-0.26(4.0)
3. PSL	0.97	0.57	-1.55 (3.4)	0*

C Adjustment matrix and diagnostics

	Adjustment matrix, L			
	Lagged Disequilibria L		Diagnostics	
	2.	3.	R^2	BP(4)[2]
1.NSD	0.33(2.4)	0.46(4.0)	0.47	0.9
2.CS	-0.13(1.3)	0*	0.46	0.8
3.PSL	-0.21(2.9)	-0.46(4.0)	0.51	1.0

(1) Asymptotic t-statistics in parenthesis. Imposed coefficients are starred.

(2) BP(k) is the Box-Pierce statistic for serial correlation of order 1 to k. Under the null of no serial correlation it is asymptotically distributed as central chi-squared with k degrees of freedom. Critical value at 5 per cent significance level $\chi^2_5(4) = 9.5$.

Table 3A: Lower level: long-run elasticities[(1),(2)]

	R_1	R_2	R_3
1. NSD	2.4	-2.8	0.7
2. CS		4.9	-2.3
3. PSL			2.7

(1) The first 3 columns show the effect of a one percentage point change in the annual rate of return R_i (i=1,2,3) on the percentage change in asset holdings (ie $(\Delta a_i/a_i)$*100).

(2) All wealth elasticities are unity.

Table 3B: Lower level: long-run impact on asset holdings (£ millions)[(1)]

	R_1	R_2	R_3	$RW^{(2)}$	$MF^{(3)}$
1.NSD	162	-202	40	33	192
2.CS	-202	378	-176	38	622
3.PSL	40	-176	136	27	290

(1) The first 3 columns show the effect of a one percentage point change in the annual rate of return R_i (i = 1,2,3) on the holdings of the i^{th} asset Δa_i (£mn).

(2) Impact on asset holdings (£mn) of £100 millon increase in wealth.

(3) The quarterly mean flow (£mn) into asset i over the period 1982-86.

Table 4: Further Tests

	Adjusted Wald Statistic, WA	Critical Value
1. 1st difference model	11.2	7.8
2. Weak separability	8.6	9.5
3. Parameter Stability 86(1)-86(4)	15.4	15.5
79(1)-79(4)	21.8	15.5

4. Co-integration test on long-run parameters in:

	DF	
	Θ_i	t_Θ
NSD	-0.46	-2.8
CS	-0.14	-2.4
PSL	-0.36	-2.5

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