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Company Sector Liquid Asset Holdings: A Systems Approach

SUCCESS IN IMPLEMENTING CONTROL of monetary aggregates requires stable asset demand functions. In the United Kingdom, aggregate studies established a stable demand for narrow money, M1 (for example, Hendry 1979, 1985), over the 1970–1984 period, but this has proved more problematic when incorporating data for the late 1980s (Hall, Henry, and Wilcox 1990). There have been even more acute difficulties in finding a stable demand function for U.K. broad money for the private sector (see Goodhart 1989 for a survey). One possible explanation for such instability is aggregation over heterogenous agents (Bank of England 1986). We seek to throw some light in this issue by examining the asset demands of the U.K. company sector. The aim of this paper is to assess how far a particular theoretical model based on the "money in the utility function approach" can explain the behavior of liquid asset holdings of the U.K. company sector, using quarterly data. There is a voluminous applied literature on the demand for financial assets of the nonbank private sector (for example, Laidler 1977, 1980; Serletis and Robb 1986; Ewis and Fisher 1984; Perraudin 1987; Backus et al. 1980; Feige and Pearce 1987; Christensen, Jorgenson, and Lau 1975; Rose 1985) but applied work, in our view, has largely neglected a systems approach to the asset decisions of the company sector (but see Jackson 1984).

Long-run asset demands are based on the AIDS model (Deaton and Muellbauer 1980) and are estimated using cointegration techniques, while in the "second stage"

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we apply the "general to specific" methodology in a systems framework. The methodology adopted allows one to "search over" alternative short-run specifications independently of the theoretically acceptable long-run cointegrating relationships. In particular our dynamic AIDS model allows one to test the theoretical restrictions of homogeneity, symmetry, and negativity which must hold if the behavior of the representative agent is to conform with the basic axioms of rational choice (for example, transitivity). We find that with a suitably flexible dynamic structure we obtain demand functions for company sector short-term assets that satisfy the theoretical restrictions implied by the AIDS model, are intuitively plausible, and exhibit parameter stability. The rest of this paper is organized as follows. In section 1 we outline the theoretical model and in section 2 we consider the modeling of short-run dynamics in a systems framework and associated econometric problems. In section 3 we discuss data problems and in section 4 we present our empirical results. We conclude with a brief summary.

1. THE AIDS MODEL

We assume multistage budgeting within the firm. "Higher level" decisions concerning employment, inventories, investment, dividends, etc. are taken independently of decisions about the distribution amongst liquid assets. This assumption of weak separability appears reasonable in the U.K. context where financial portfolio decisions are largely managed by Corporate Treasurers. It also makes the model tractable.

The representative agent is assumed to distribute wealth among alternative assets in order to minimize the cost of achieving a given level of utility.[1] The axioms of rational choice in demand theory (that is, 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 select the PIGLOG (Price Independent Generalized Logarithmic) which, in common with others [for example, indirect translog (Christensen et al. 1975)] is a second-order approximation. Within the PIGLOG class we choose the AIDS cost function (Deaton and Muellbauer 1989).

The budget constraint is

$$\sum_i p_{it} \, a^T_{it+1} = W^T_t \tag{1}$$

[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. Hence asset holdings represent purchasing power over future consumption goods. Friedman (1956) discusses the theory of asset demands based on utility and Barnett (1980) develops these ideas in a more rigorous framework.

[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.

where $p_{it} = [(1 + r_{it})(1 - g_t)]^{-1}$; r_{it} = *expected* (proportionate) nominal return on asset i, between t and $t + 1$ (including any capital gains); g_t = *expected* (proportionate) rate of goods price inflation between t and $t + 1$; a^{τ}_{it+1} = real asset holdings at end of period $t + 1$, $(= a_{it+1}/z_{t+1})$; a_{it} = nominal asset holdings of the ith asset at end of period t; Z_t = goods price index; W^{τ}_t = real wealth at end of period t, $(= W_t/Z_t)$.

Solving the constrained cost minimization problem leads to the AIDS share equations (see, for example, Barr and Cuthbertson 1991; Weale 1986):

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

where $s_i = (a_{it}/W_t)$, $\ln P^*_t = \Sigma\, \bar{s}_i \ln p_{it}$. Note that $\ln P^*_t$ may be interpreted as a composite real discount factor (s_i are the sample mean shares). The AIDS share equations are linear in the parameters whereas the share equations from the indirect translog utility function are nonlinear. The AIDS model is therefore far more tractable when we seek to incorporate and interpret the coefficients of an interdependent dynamic adjustment process (see section 2). The theoretical restrictions of the AIDS model are as follows. The adding up constraints:

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

Homogeneity:

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

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

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

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

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

is negative semidefinite (δ_{ij} is the Kronecker delta).

Thus our systems approach implicitly imposes data admissibility 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. One might also wish to judge the model on more intuitive notions, (for example, that own price effects are negative, wealth elasticities are "reasonable," etc).

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

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

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

The semielasticities of asset holdings with respect to the *annual percentage* rate of return R (that is, $r = R/400$) are

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

2. DYNAMIC ADJUSTMENT AND ECONOMETRIC ISSUES

The share equations (2) from the constrained maximization problem we designate as *desired* long-run shares $\mathbf{s}_t^*$ which may be represented in vector notation:

$$\mathbf{s}_t^* = \mathbf{\Pi X}_t \tag{5}$$

$\mathbf{s}_t^* = k \times 1$ vector of desired long-run asset shares, $\mathbf{X}_t = q \times 1$ vector of independent variables, $\mathbf{\Pi} = k \times q$ matrix of long-run parameters. The adjustment of actual shares $\mathbf{s}_t$ to desired long-run shares is assumed to operate via a generalized error feedback mechanism (Brainard and Tobin 1968; Smith 1975; Anderson and Blundell 1983):

$$\Delta \mathbf{s}_t = \mathbf{\Pi}^* \Delta \mathbf{X}_t + \mathbf{L}\,(\mathbf{s} - \mathbf{s}^*)_{t-1} + \boldsymbol{\varepsilon}_t \tag{6}$$

where the disequilibria in $(k - 1)$ asset shares at time $t - 1$ influence the current period adjustment of any particular asset share. Since $\Sigma_1^k \, (s_i - s_i^*)_{t-1} = 0$, only $(k - 1)$ independent disequilibrium shares are required in (7) [that is, $\mathbf{L}$ is $(k \times (k - 1)]$. Adding-up restrictions imply that the columns of $\mathbf{\Pi}^*$ and $\mathbf{L}$ sum to zero. The system is dynamically stable if the eigenvalues of the appropriate adjustment matrix have modulus less than unity.[3]

Prior to the use of cointegration techniques estimation of (6) would have proceeded by running the unrestricted set of equations ($\mathbf{u}_t$ is a white noise error term):

$$\mathbf{s}_t = \mathbf{R}_1 \, \mathbf{s}_{t-1} + \mathbf{R}_2 \, \mathbf{X}_t + \mathbf{R}_3 \, \mathbf{X}_{t-1} + \mathbf{u}_t \; . \tag{7}$$

[3]If the disequilibrium term for asset 1 is excluded, then the *estimated* adjustment matrix is $\mathbf{L} = (\mathbf{l}_2, \mathbf{l}_3, \ldots \mathbf{l}_k)$ where $\mathbf{L}$ is kx$(k - 1)$. The dynamics of the full model may be written $\mathbf{s}_t = (\mathbf{I}_k + \mathbf{L}^*)\mathbf{s}_{t-1}$, where, $\mathbf{s}_t$ is $(k \times 1)$, $\mathbf{L}^* = (\mathbf{0}, \mathbf{l}_2, \mathbf{l}_3, \ldots \mathbf{l}_k)$ is $(k \times k)$, $\mathbf{0} = (0, 0 \ldots 0)$ is $(k \times 1)$, and $\mathbf{I}_k = (k \times k)$ identity matrix. One of the eigenvalues of $(\mathbf{I}_k + \mathbf{L}^*)$ is unity and stability requires that the other $(k - 1)$ eigenvalues have negative real parts.

The main disadvantages of this approach are as follows. First, in testing down to a parsimonious dynamic representation (via restrictions on the R_i matrix elements) one implicitly alters the long-run solution, and the final equation (possibly after considerable "search-time" has been invested) may be unacceptable on a priori grounds. Second, we cannot be sure that the ensuing long-run solution yields a cointegrating vector or that there is a cointegrating vector among the set of variables included in (6). Third, in the absence of cointegration the usual test statistics in (6) do not apply (Engle and Granger 1987). Fourth, long-run theory restrictions (for example, symmetry and homogeneity) are often not imposed (for instance, Weale 1986) although procedures are available (see Bewley 1979). For the above reasons we think it worth exploring a systems approach using cointegration techniques.

Cointegration establishes a parameter vector which yields stationary I(0) errors (Granger 1986; Engle and Granger 1987). Assuming all variables in the long-run share equations (5) are integrated of order one I(1), and yield a cointegrating vector, then in the "first-stage" regression, OLS on (5) (using actual shares $\mathbf{s}_t$) yields super-consistent estimates, of $\mathbf{\Pi}$ (Stock 1984). The residuals from $(k - 1)$ of the cointegration equations $(s_i - \hat{s}_i)$ are then substituted into the dynamic system error-feedback equations (6) (Hall 1986). In equation (6), $\Delta \mathbf{s}_t$ and $\Delta \mathbf{X}_t$ are stationary I(0) variables by construction while the individual elements of the vector $(\mathbf{s} - \mathbf{s}^*)_{t-1}$ are also I(0) because we have established cointegrating vector $\mathbf{\Pi}$ in the stage-one regressions. Hence by the Granger representation theorem (Engle and Granger 1987) there exists an error-correction representation of the form (6) and the usual test statistics apply (Engle and Granger 1987). The "general-to-specific" methodology (Hendry 1989) may be applied in the "second stage" to obtain parsimonious dynamic equations while holding the long-run parameters fixed (Hendry, Pagan, and Sargan 1984). Although attractive, there are some practical problems with the two-step procedure. The cointegration regression estimates may suffer from small sample bias (Hendry 1986) and the cointegrating vector may not be unique. However, given the relatively strong theoretical restrictions to be placed on the long-run share equations (for example, homogeneity and symmetry, negative own "price" effects), we are mainly interested in finding a set of plausible parameter estimates that conform to theory and form a cointegrating vector. We are therefore willing to risk some small sample bias at stage one (and possibly an inferior fit of the final equation) in order to obtain a theoretically consistent approach. We therefore adopt an informal approach, trading off fit in the second-stage regressions against the system restrictions implied by our theoretical model. The final parsimonious system of equations is subject to the usual test procedures (Hendry 1985).

In order to impose cross-equation restrictions on the long-run parameters we use maximum likelihood with a diagonal covariance matrix (obtained form running OLS on each equation separately). When estimating the dynamic short-run equations we report results using 3SLS (Zellner and Theil 1962). 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.

Restrictions on parameters are tested using Wald tests and a quasi-likelihood ratio test statistic QLR. When using IV the usual likelihood ratio test (Gallant and Jorgenson 1979) is $QLR = T(Q_0 - Q_1)$, where Q_0 (Q_1) is the value of the minimum distance criterion for the null (maintained) hypothesis; $Q = \mathbf{e}' (\boldsymbol{\Sigma}^{-1} \otimes \mathbf{N})\, \mathbf{e}$, where $\mathbf{b}$ = vector of parameters, $\mathbf{e}$ = a stacked vector of residuals, with variance-covariance matrix $\boldsymbol{\Sigma}$ and projection matrix of instruments $\mathbf{N}$). When using QLR, the same set of instruments must be used under both H_0 and H_1 and $\boldsymbol{\Sigma}$ must be held fixed at its value under the maintained hypothesis.

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 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.

3. DATA USED

We use quarterly data and the asset categories modeled are the following "short-term" assets of the company sector: M1 = transactions balances (= notes and coin + sterling sight deposits); TD = time deposits (= sterling time deposits + building society deposits + local authority temporary debt); PSD = public sector debt (British Government Securities, Northern Ireland Central Government Long-Term Debt and Treasury Bills); FCD = foreign currency deposits.

In order to apply the above model to the company sector we delineate the problem by assuming decisions concerning the portfolio of short-term assets are weakly separable from other asset decisions (and real decisions). We also assume weak intertemporal separability. Evidence in Mayer (1988) and Chowdhury, Green, and Miles (1987) suggests that our asset split may be a reasonable working hypothesis.

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 *all* elements in the matrix satisfy the accounting identities (that is, 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. Data on the sight-time deposit split is only available from 1975 (III).

The mean shares for our asset categories over the period 1977–86 are M1 (23 percent), TD (55 percent), PSD (5 percent), FCD (17 percent). However, there has been a relatively large secular fall in TD from 63 percent in 1977 to around 48 percent in 1986 and this mirrored by the rapid rise in FCD after 1979 from a share of 8 percent to 23 percent by 1983/4. Tax instruments were initially included in the asset set but proved problematic, probably because of the complexity of the "true" rate of return (see Jackson 1984). The rate of return on M1 is the negative of the expected inflation rate (of the price of total final expenditure, TFE) and for TD the own rate is taken to be the rate on "parallel money market" assets, namely, the three-month rate on Local Authority bills. Various rates of return for PSD and FCD were tried, all of which included the running yield and capital gains. The running yield plus a three-year backward-looking capital gain is used for the return on PSD and FCDs. The F-T actuaries price index for all government stock and the running yield are used for PSD while the return on FCD is the yield on three-month dollar deposits in London plus capital gains due to changes in the dollar-sterling exchange rate. (One-quarter-ahead and one-year-ahead capital gains variables were also tried but gave very unsatisfactory results.) All rates of return are net of the marginal corporate tax rate.

The "static" cointegrating equations can be estimated by OLS (Stock 1987) but the second-stage regressions are estimated by 3SLS (see below) using "errors in variables" for expectations variables (for example, inflation) and instruments for current-period AIDS prices and wealth. All data used are seasonally unadjusted but seasonal dummy coefficients are not reported. The regressions are run over the period 1976.IV–1986.IV. Critical values of test statistics are given at a 5 percent significance level (unless stated otherwise).

4. EMPIRICAL RESULTS

The long-run AIDS share equations are given in equation (2) and we first establish the order of integration of the variables. Phillips and Perron (1988) provide tests for unit roots which are robust to a wide variety of serial correlation and time-dependent heteroskedasticity and also allow for the possibility of a deterministic as well as a stochastic trend in the data.[4]

The Phillips-Perron tests involve OLS regressions of the form

$$y_t = \mu^* + \alpha^* y_{t-1} + u_t^* \tag{8}$$

$$y_t = \tilde{\mu} + \tilde{\beta}(t - n/2) + \tilde{\alpha} y_{t-1} + \tilde{u}_t \tag{9}$$

where the error terms are stationary ARMA processes (possibly) with time-dependent variances (t = time trend and n = number of observations). The null hypotheses of a unit root, with or without drift, that is, $H_0^1{:}\alpha^* = 1$ and $H_0^2{:}\mu^* = 0$, $\alpha^* = 1$, are tested against the stationary alternatives by means of the adjusted t- and

[4]Schwert (1987, 1988) notes some deficiencies in both the Dickey-Fuller (Dickey-Fuller 1979) and Phillips-Perron tests in the presence of large positive moving average parameters. Perron (1989) notes the difficulties of distinguishing between (i) deterministic trend process with a regime shift and (ii) a stochastic trend. Strong a priori assumptions concerning the timing of the regime shift are required. Although the shares s_i are I(1) in the data set considered, there is a theoretical problem in that shares are bounded. The tests used are for a random walk which is a particular case of a nonstationary I(1) series. Clearly shares cannot follow a random walk (which is unbounded) but they may be nonstationary. Similar practical difficulties arise with variables such as bilateral exchange rates (which are bounded below) and the percentage unemployed, when applying unit root and cointegration tests. We are grateful to the referees for helping to clarify our views on this issue.

TABLE 1

PHILLIPS-PERRON TESTS FOR UNIT ROOTS

	$Z(t\alpha^*)$	$Z(\Phi_1)$	$Z(t\tilde{\alpha})$	$Z(\Phi_3)$	$Z(\Phi_2)$
s_1 (M1)	−1.8	2.7	−2.1	2.5	1.6
s_2 (TD)	−1.8	3.0	−2.6	3.9	3.1
s_3 (PSD)	−2.0	3.5	−2.5	4.6	3.1
s_4 (FCD)	−1.1	1.1	−1.4	1.1	1.5
ln (W^T/P^*)	0.9	2.1	−1.1	1.9	2.5
$\ln p_1$	−2.2	2.7	−3.0	5.5	3.7
$\ln p_2$	−2.1	3.0	−3.0	6.2	4.2
$\ln p_3$	−2.2	3.5	−2.8	4.2	3.6
$\ln p_4$	−1.5	1.1	−1.3	1.1	0.7
Δs_1	−8.6	37.0	−8.5	36.2	24.2
Δs_2	−7.2	25.5	−7.2	26.0	17.3
Δs_3	−6.4	20.8	−6.7	22.3	14.9
Δs_4	−6.6	21.4	−6.6	21.4	14.3
$\Delta \ln (W^T/P^*)$	−5.2	13.4	−5.5	15.1	10.1
$\Delta \ln p_1$	−7.5	27.6	−7.3	26.7	17.8
$\Delta \ln p_2$	−8.5	35.4	−8.3	33.9	22.7
$\Delta \ln p_3$	−8.6	36.9	−8.6	36.9	24.6
$\Delta \ln p_4$	−6.8	22.9	−6.7	22.4	15.0

NOTE: The critical values at 5 percent significance level (see Dickey and Fuller 1979) for the test statistics in columns 1 to 5 are −2.86, 4.86, −3.4, 6.73, 5.13, respectively. If the absolute value of the test statistic exceeds its critical value then we reject the null hypothesis of a unit root.

F-statistics $Z(t\alpha^*)$ and $Z(\Phi_1)$ (Table 1). Equation (9) allows for a deterministic trend and the null hypotheses, $H_0^3{:}\tilde{\alpha} = 1$; $H_0^4{:}\tilde{\beta} = 0, \tilde{\alpha} = 1$; and $H_0^5{:}\tilde{\mu} = \tilde{\beta} = 0, \tilde{\alpha} = 1$ are tested using $Z(t_{\tilde{\alpha}})$, $Z(\Phi_3)$, and $Z(\Phi_2)$ respectively.

With equation (8) or (9) as the maintained hypotheses we cannot reject the null hypothesis that all the data series in Table 1 for the levels of the variables contain a unit root [that is, are at least I(1)][5] at a 5 percent significance level. However, the test statistics reported in the lower half of Table 1 imply that we easily reject the null hypothesis that the first-difference in the variables are nonstationary. We take these results to indicate I(1) variables.

Below, we first present our preferred equations and their economic interpretation. We then discuss general-to-specific tests of the model, tests of parameter stability, and other variants considered.

Long-Run Share Equations

The preferred long-run share equations with symmetry and homogeneity imposed are shown in Table 2. The parameters of the AIDS model are not readily interpretable in terms of the usual economic concepts (for example, elasticities) but the following general features are of interest.

[5]The lag length for the order of the moving average component of the error term is taken to be 2 in Table 1. The results are very similar for lag lengths $m = 1$ to 4.

TABLE 2

LONG-RUN COEFFICIENTS: HOMOGENEITY AND SYMMETRY IMPOSED

		Price Matrix			
	g^a	$\ln p_2$	$\ln p_3$	$\ln p_4$	$\ln (W^T/P^*)$
s_1 (M1)	0.94 (1.0)	−0.87 (1.3)	−0.94 (3.0)	0.87 (3.5)	−0.05 (2.8)
s_2 (TD)	−0.87	−1.12 (1.8)	1.43 (5.6)	0.56 (1.6)	−0.15 (5.3)
s_3 (PSD)	−0.94	1.43	−0.64 (3.5)	0.15 (1.4)	−0.04 (5.4)
s_4 (FCD)	0.87	0.56	0.15	−1.58 (3.4)	0.24 (6.7)

	Θ_i^a	DF[a]	$Z(T\hat{\alpha})$	$Z(t_\alpha^*)$	BP(1)[b]	BP(4)[b]
s_1 (M1)	−0.35	−3.2	−3.2	−3.1	0.8	6.1
s_2 (TD)	−0.23	−2.7	−2.6	−2.5	6.1	9.7
s_3 (PSD)	−0.41	−3.3	−3.3	−3.3	6.2	13.7
s_4 (FCD)	−0.14	−2.1	−2.4	−2.1	4.9	12.4

NOTES: t-statistics are in parentheses but these are not distributed as a students' t-distribution. They cannot be used for hypothesis testing.
[a] Θ_i is the coefficient in the regression $\Delta e_{it} = \Theta_i e_{it-j}$, where e_i = residuals from the cointegrating regression. The $Z(t\hat{\alpha})$ and $Z(t_\alpha^*)$ statistics are the Phillips-Peron (1988) tests for a unit root in the residuals for lag length $m = 4$: the tests therefore allow for moving average errors up to order 4. $Z(t\hat{\alpha})$ is the Phillips-Perron adjusted t-statistic for a unit root in the residuals in equation (2) in the text, with $\alpha^* = 0$ imposed: $Z(t_\alpha^*)$ has $\alpha^* \neq 0$. (Results are qualitatively unchanged for lag lengths m = 1–4.) The critical values for the DF, $Z(t\hat{\alpha})$ and $Z(t_\alpha^*)$ statistics at a 10 percent significant level are approximately −4.3. These are based on simulation (Engle and Yoo 1987).
[b] BP (k) is the Box-Pierce statistic for serial correlation of order l to k. Under the null of no serial correlation it is asymptotically distributed as central chi-squared with k degrees of freedom. Critical values at 5 percent significance level are $\chi_c^2(1) = 3.8$, $\chi_c^2(4) = 9.5$.

a. The own-rate price coefficients for TD, PSD, FCD, and γ_{ii} are all negative. The γ_{ij} ($i \neq j$) for TD, PSD, and FCD are positive indicating these assets are substitutes.
b. Except for the equation for M1, the Box-Pierce, BP(1), statistics indicate severe first-order serial correlation in these "static" cointegrating regressions. Hence these t-statistics cannot be used for hypothesis testing (and merely give an indication of the relative contributions of the independent variables to movements in the dependent variable).[6]
c. The matrix of k_{ij} coefficients [see equation (3d)] evaluated at mean values of variables has all eigenvalues negative and is therefore negative semidefinite.

The results without imposing any restrictions on the Γ matrix are qualitatively similar to those reported above except for γ_{22} which is positive. The Dickey-Fuller (DF) tests for the four unrestricted equations are −5.0, −5.0, −4.1, and −4.6 and, given a critical value of −4.3 (at 10 percent significance level), strongly suggest the presence of a cointegrating vector for each equation. Imposing symmetry and homogeneity yields $\gamma_{22} < 0$ (Table 2) and therefore "improves" the cointegrating vector on a priori grounds (we test this restriction below). However, imposing these restrictions weakens the case for a set of cointegrating vectors since the DF and Phillips-Perron tests are now below their critical values (Table 2).[7] These statistics

[6]West (1988) provides a correction for the t-statistics from a single cointegrating regression but this may not be applicable here because not all of the variables have a nonzero drift parameter, and we are dealing with a system of equations.

[7]Under the null of a random walk the cointegrating Durbin Watson (CRDW) statistic is uniformly most powerful (against a stationary Markov process). The CRDW statistics for equations (1)–(4) of Table 2 are 0.35, 0.23, 0.41, and 0.14 and the critical value at a 10 percent significance level is about 0.35 (Engle and Yoo 1987). The latter results are indicative of stationary residuals (but not definitive) as are the correlograms which all have (absolute) values below 0.1 at lag lengths 6 to 16. Lack of a definitive

TABLE 3

LONG-RUN ELASTICITIES

	g^a	R_2	R_3	R_4	Wealth[†]
1. M1	0.98	0.77	0.97	−0.95	0.78
2. TD	−0.39	0.63	−0.67	−0.30	0.72
3. PSD	−4.7	−7.2	3.40	−0.79	0.20
4. FCD	1.3	−1.0	−0.23	2.5	2.4

NOTES: The first four columns show the effect of a one percentage point change in g^a (the *annual* inflation rate), and *annual* rates of return R_i (i = 2, 3, 4) on the percentage change in asset holdings (that is, $(\Delta a_i/a_i)$*100).
[†]Wealth elasticity [that is, $(\Delta a_i/a_i)\,(\Delta W/W)^{-1}$].

have low power against highly dynamic stationary alternatives (Hendry 1986; Hendry and Ericsonn 1988) and we therefore use these statistics as a "useful guide to decide when to impose the cointegration constraint" (Engle and Yoo 1987, p. 159). On balance, we therefore proceed under the assumption that the model in Table 2 yields a set of cointegrating vectors that are also acceptable on a priori grounds.[8]

Long-Run Elasticities and Semielasticities

The semielasticities with respect to rates of return and the wealth elasticities are shown in Table 3. The own rate semielasticities of TD, PSD, and FCD are 0.63, 3.4, and 2.5, respectively. At higher rates of inflation there is a net move out of total domestic assets and into FCD. Higher inflation may imply a lower expected sterling exchange rate and hence a higher expected return on FCD which may not have been picked up by our backward-looking price variable for FCD. In terms of the change in the absolute level of asset holdings a major switch at high rates of inflation is out of PSD and into FCD (Table 4, column 1): there is hardly any net switching out of total domestic capital certain assets M1 and TD (although the positive coefficient on M1 is somewhat counterintuitive). A rise in the return on TD leads to a switch out of PSD and FCD and into M1 and TD (Table 4, column 2). PSD and M1 are comple-

conclusion using unit root tests in finite samples often occurs in applied work [see, for example, Hendry and Ericsonn (1988) and Blanchard (1989, p. 1151) who states that "results . . . must be seen as dependent on a priori assumptions on the time series properties of the series"]. Note that the existence of a dynamic error correction model requires cointegration between the levels of I(1) variables and vice versa (Engle and Granger 1987). Hence the statistical acceptability of the final dynamic error correction model is also an indirect test of cointegration. Ultimately one's choice of model is an amalgam of a priori views and statistical evidence from the cointegrating regression and the dynamic error correction model.

[8]Further evidence on the existence of a cointegrating set of variables may be obtained from Johansen's (1988) maximum likelihood estimation of the cointegrating vector. The Johansen procedure also provides a method of determining the number of unique cointegrating vectors **r** in a single equation context although these tests have not been developed for a set of equations subject to cross-equation restrictions. However, we did apply the Johansen procedure to our *unrestricted* cointegrating equations. For all four equations we decisively reject the hypothesis of no cointegrating vectors (**r** = 0) on a likelihood ratio test. The likelihood ratio tests for each of the four equations (with 5 percent critical value of 88.1) are 108.4, 105.8, 109.1, and 112.7, respectively. The test for a unique cointegrating vector (**r** < 1) was just acceptable for two equations and just rejected for the remaining two. The likelihood ratio statistics for **r** < 1 for equations (1)–(4) were 69.1, 70.2, 66.5, and 72.3 respectively (with 5 percent critical value of 70.0). Hence we can be fairly confident that we have at least one cointegrating vector for all four equations but for equations (2) and (4) (in Table 2) the cointegrating vector may not be unique.

TABLE 4

LONG-RUN IMPACT ON ASSET HOLDINGS (£M)

	g^a	R_2	R_3	R_4	Wealth†	MF††
1. M1	63	59	63	−59	19	270
2. TD	−59	75	−96	−38	39	450
3. PSD	−63	−96	43	−10	2	24
4. FCD	59	−38	−10	107	40	278

NOTES: The first four columns show the effect of a one percentage point change in g^a (the annual inflation rate) and annual rates of return R_i (1 = 2, 3, 4) on the holdings of the ith asset Δa_i (£m).
†Impact on asset holdings Δa_i (£m) of £100m increase in wealth. To satisfy the budget constraint ($\Sigma a_i = W$) this column sums to 100.
††The quarterly mean flow (£m) into asset i over the period 1982–86.

ments (Table 4, column 3) while M1 and TD are relatively strong substitutes with FCD (Table 4, column 4).

All assets have positive wealth elasticities (Table 3, final column). An increase in the level of wealth of £100 m leads to an increase in M1 of £19m, TD by £39m, PSD by £2m, and FCD by £40m [based on mean shares; see equation (4a)]. The high wealth elasticity for FCD of 2.4 reflects in part the increasing attractiveness of such assets after the ending of exchange controls in 1979.

Short-Run Equations (Table 5)

The short-run equations are estimated by 3SLS treating the AIDS prices and wealth as endogenous. The instruments used are two lagged values ($t - 1, t - 2$) of all AIDS prices, wealth, and real total final expenditure and two lags of the three-month Eurodollar rate. Drawing on (2) and (6), the short-run equations may be represented as follows:

$$\Delta \mathbf{s}_t = \mathbf{C}\Delta \ln \mathbf{p}_t + \mathbf{K}\Delta \ln(W^T/P^*)_t + \mathbf{L}(\mathbf{s} - \mathbf{s}^*)_{t-1} \qquad (10)$$

TABLE 5

SHORT-RUN PRICE, AND ADJUSTMENT COEFFICIENTS, HOMOGENEITY AND SYMMETRY IMPOSED

5A. Price Matrix **C**

	Δg	$\Delta \ln p_2$	$\Delta \ln p_3$	$\Delta \ln p_4$
Δs_1 (MI)	0*	0*	0*	0*
Δs_2 (TD)	0*	−1.18 (2.1)	0.18 (0.9)	1.0 (2.1)
Δs_3 (PSL)	0*	0.18	−0.45 (1.7)	0.27 (1.1)
Δs_4 (FCD)	0*	1.0	0.27	−1.27 (2.5)

5B. Adjustment Matrix **L**

	Lagged Disequilibria				Diagnostics	
	$(s_2 - s_2^*)_{t-1}$	$(s_3 - s_3^*)_{t-1}$	$(s_4 - s_4^*)_{t-1}$	R^2	BP(1)†	BP(8)†
Δs_1 (MI)	0.81 (4.5)	0.79 (2.2)	0.52 (2.8)	0.25	0.1	7.5
Δs_2 (TD)	−0.81 (4.5)	−0.93 (2.3)	−0.26 (1.3)	0.54	0.3	15.7
Δs_3 (PSL)	0*	−0.45 (2.2)	0*	0.36	0.2	11.2
Δs_4 (FCD)	0*	0.59 (1.6)	−0.26 (1.8)	0.37	0.1	15.3

NOTES: An asterisk * indicates the coefficient has been constrained.
† 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 values at 5 percent significance level are $\chi^2_c(1) = 3.8$, $\chi^2_c(8) = 15.5$.

where **C**, **K**, and **L** are suitably dimensioned matrices of short-run parameters.

The main results (see Table 5) are as follows: Homogeneity and symmetry on the short-run price matrix (**C**) are imposed (Table 5A), and are not rejected (at a 5 percent significance level) on a Wald test ($W(3) = 3.9$, $\chi_c^2 = 7.8$; $W(6) = 13.5$, $\chi_c^2 = 21.0$). All own-rate coefficients are correctly signed (that is, negative) and larger than the cross-rate effects. In the short run, TD, PSD, and FCD are substitutes.

Statistically, the first row and column of the **C** matrix could be constrained to zero. (The Wald test is $W(3) = 0.8$, $\chi_c^2 = 7.8$.) This implies no short-run impact of prices on M1 and no independent short-run effect on asset shares from changes in inflation. The short-run wealth effects can be constrained to zero ($W(3) = 1.2$) implying a short-run wealth elasticity of unity for all assets.

The adjustment matrix **L** is shown in Table 5B. All diagonal elements are negative indicating that excess holdings of asset i leads to a fall in the share of asset i in the subsequent period. The three zero restrictions on the **L** matrix are not rejected ($W(3) = 1.7$, $\chi_c^2 = 7.8$). The eigenvalues of the (augmented) **L** matrix have "real parts" -0.19, 0.55, 0.74, indicating a convergent response after a one-quarter lag.

Considering that the equations are explaining changes in asset shares the R^2 (Table 5B) are reasonable. The Box-Pierce statistics indicate that there is not an acute problem of serial correlation (although the equation for TD may have some negative second-order serial correlation). The interdependent error feedback formulation therefore appears to provide a data coherent (Hendry, Pagan, and Sargan 1984) dynamic model.

Tests of *long-run* symmetry, homogeneity and homotheticity using only the first-stage cointegrating regressions are not yet available for a system of equations [but see Johansen (1988) and West (1988) in a single-equation framework] as the usual t and F-statistics do not apply. We therefore tested these long-run parameter restrictions in a "general" unrestricted ADL model (see Engle and Granger 1987) estimated by 3SLS [see equation (7)]. The Wald test for long-run homotheticity [that is, $\beta_i = 0$, ($i = 1, 2, 3$); see equation (2)] indicates rejection: $W(3) = 9.0$, ($\chi_c^2 = 7.8$). Long-run symmetry (and homogeneity) is not rejected: $W(12) = 5.9$, ($\chi_c^2 = 12.6$). The test for long-run symmetry (and homogeneity), and the other restrictions on the parameters in Table 2 are accepted on a Wald test ($W(15) = 6.3$, $\chi_c^2 = 25$) and on a QLR test ($QLR(15) = 7.6$).

Dropping the last four and eight data points yields no qualitative changes in the short-run parameters and the equation easily passes (our Wald systems analogue to) the Salkever test for the 1985.I–1985.IV and 1985.I–1986.IV periods ($W(12) = 6.1$, $W(24) = 23.0$, respectively, $\chi_c^2(12) = 21.0$, $\chi_c^2(24) = 36.4$), indicating parameter stability. (There is only one t-statistic on the dummy variables in excess of 1.1, namely, for 1986.II in the TD equation, $t = 2.5$). We can test our EFE against a first-difference model by testing that all the (independent) elements of the adjustment matrix **L** [equation (11)] are zero: the first-difference model is easily rejected ($W(9) = 56.6$, $\chi_c^2 = 16.9$). The static model is rejected since it has severe first-order serial correlation (see BP(1) statistics, Table 2) which is indicative of misspecified dynamics.[9]

5. SUMMARY

Within a systems framework we find that the company sector's demand for liquid assets depends on movements in relative prices, and these responses are plausible on a priori grounds. Our asset demand system is estimated using the Granger-Engle two-step procedure. This has the advantage of enabling a general-to-specific search over the short-run dynamics after establishing theoretically acceptable long-run cointegrating equations. In a system with a large number of potential parameters to be estimated, this is a major practical advantage. However, one cannot "get something for nothing" and in small samples the cointegrating vector may be biased (and not unique). The success of this approach is to be judged partly against alternative systems modeling procedures (see, for example, Bewley 1979; Anderson and Blundell 1983; Owen 1986; Barr and Cuthbertson 1989) and on how well the system fits the data and conforms to theoretical priors. While the fit is no doubt not as good as could be found by an equation-by-equation general-to-specific search procedure, nevertheless, this is to be weighed against the insights obtained by considering a coherent theoretical model based on the axioms of rational choice. Given the limited data set available we feel the methodology has yielded reasonable results and that the method could be usefully applied in modeling other sectors and systems of equations.

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[9]Although tests are biased because of the presence of serial correlation, nevertheless, homogeneity ($W(3) = 27.8$, $\chi^2_c = 7.8$) and symmetry ($W(6) = 48.6$, $\chi^2_c = 12.6$) are decisively rejected, in the static model.

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