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The Demand for M1 in the U.S.A., 1960-1988

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Estimated U.S. M1 demand functions appear unstable, regularly "breaking down," over 1960-1988 (e.g. *missing money, great velocity decline, M1-explosion*). We propose a money demand function whose arguments include inflation, real income, long-term bond yield and risk, T-bill interest rates, and learning curve weighted yields on newly introduced instruments in M1 and non-transactions M2. The model is estimated in dynamic error-correction form; it is constant and, with an equation standard error of 0·4%, variance-dominates most previous models. Estimating alternative specifications explains earlier "breakdowns," showing the model's distinctive features to be important in accounting for the data.

I. INTRODUCTION

In the early 1970's, the demand function for money in the United States (as measured by M1), seemed to be well established as a relatively simple and constant empirical equation with a sound theoretical foundation. Three subsequent episodes of breakdown in this relationship, namely 1974(1)-1976(2), 1982(1)-1983(2) and 1985(1)-1986(4), have strained such a view and could be interpreted to suggest that the demand *function* may be subject to abrupt shifts. In each of these periods, actual money-holding behaviour deviated dramatically from the predictions of existing models.[1] There is no shortage of explanations for the apparent change in underlying behaviour, including the availability of new financial instruments or new techniques of financial management, but many of the explanations have not been fully susceptible to quantitative expression and testing. Further, any explanation, to avoid the charge that it represents *ad hoc* rationalization, should be equally capable of explaining the apparent downward shift in money demand of the mid-70's and the apparent upward shifts of the 1980's.

1. The mid-70's period is known as the *Case of the Missing Money*, when the quantity of money held fell dramatically below model predictions (see Goldfeld (1973), (1976)). The early 80's period, when money held substantially exceeded most projections, is denoted the *Great Velocity Decline* by the Federal Reserve Bank of San Francisco (1983). We refer to the rapid growth of real M1 by more than 20% over 1985(1)-1986(4) as the *M1-explosion*, following Hendry and Ericsson (1991).

Estimated money demand *equations* displaying shifts may simply be mis-specified (see, for example, econometric evidence for the U.K. in Hendry (1979)). In periods of relative tranquility or of slow change in financial markets and institutions, model mis-specification may not be apparent. During more volatile periods, mis-specified models will be discovered to "shift" or demonstrate predictive failure. Meanwhile, a correctly specified model may show no statistically significant structural change. From this perspective, a progressive research strategy is called for: any new model must not only be constant, it must account for previous failures and successes. Surveying the literature (for reviews see Judd and Scadding (1982), Laidler (1985), Cuthbertson (1989) and Goldfeld and Sichel (1990)), a number of potentially important features remain to be incorporated jointly in U.S. M1 demand models.

The standard demand for money function is of the form (see Goldfeld (1973)):

$$\ln M/P = \alpha + \beta_1 \ln y + \beta_2 \ln R_s + \beta_3 \ln R_2 + \beta_4 \ln (M/P)_{-1} \tag{1}$$

where ln denotes the natural logarithm, α, $\beta_1 > 0$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$ are assumed constant, and M/P, y, R_s and R_2 respectively denote real M1, real GNP, a short-term market interest rate, and an interest rate on interest-bearing deposits in M2 (typically a passbook rate); all variables are measured quarterly and $(M/P)_{-1}$ is the one-period lagged dependent variable. This function is potentially mis-specified in four ways: incomplete dynamic structure; inadequate treatment of M1's own yields and those of alternative monetary instruments in the face of financial innovation; inappropriate treatment of inflation (by its complete exclusion); and omission of yield and risk of other assets (e.g. long-term bonds). We propose a model reformulated to allow resolution of such a claim. The result is an M1 demand model that passes tests of parameter constancy over all the challenging sub-periods of the twenty-eight years 1960–1988, including the three intervals identified above as problem areas for other demand equations. Experiments with the model, omitting distinctive variables and dynamics, result in parameter constancy rejection and errors of fit comparable to those experienced by earlier alternative models. We infer that the reason for the shifts in alternative models is their omission of appropriate dynamic structure and of important variables.

The dynamic structure of our equation is based on an error-correction model, with contemporaneous and lagged conditioning variables (see Sargan (1964), Davidson *et al.* (1978) and Hendry, Pagan and Sargan (1984)). Interest rate variables are designed so that the introduction of new financial instruments in M2 does not alter the model's structure, with returns to newly introduced assets being adjusted for learning. Inflation is separately included. It was identified by Cagan (1956) and Friedman (1956) as a primary determinant of money demand during hyper-inflation. As with M1 models in the U.K., inflation turns out to be an important determinant of U.S. money demand at moderate inflation rates as well, especially in accounting for the great velocity decline, 1982–1983 (as noted by Rose (1985) and Gordon (1984)). Finally, Tobin (1958) argued that the risk of long-term bond holding should influence M2 demand. We formalize the risk variable in a specific way and find a strong impact as well on M1 demand; bond yield and holding-period risk provide an important part of the explanation for the episodes of model failure. Such episodes are both a problem and an opportunity: explanations are badly needed, and the great variation in experience provides an unusually discriminating test of alternative hypotheses about money-demand behaviour.

Section II discusses the background economic model and Section III the econometric approach. The specification is discussed in Section IV and its implementation in Section V. Section VI notes alternative specifications and relates this study to earlier research on

U.S. M1, and Section VII provides a detailed analysis of the historical episodes which were problematic for previous models. Section VIII concludes, with an appendix on the data series employed.

II. FINANCIAL RISK AND INFLATION IN MONEY DEMAND

It has long been recognized that the demand for a safe asset, e.g. money or Treasury bills, depends on a risk-return tradeoff against higher yielding risky assets (see e.g. Tobin (1958)). However, in a model that includes both money as a low-yielding transactions instrument and bills as a higher-yielding investment instrument, it is usually argued that bills dominate money as a portfolio asset (see e.g. Ando and Shell (1975), Chang, Hamberg and Hirata (1983) and Goldfeld and Sichel (1990)). Hence it is argued, in particular by Ando and Shell, that the risk-return tradeoff determines the portfolio balance between bills and risky assets, and money demand is determined by a yield-transaction cost tradeoff without allowance for risk. We argue, on the contrary, below that there may be a relevant risk-return tradeoff helping to determine the demand for money, even in the presence of safe higher-yielding assets. The basis of this contention is a capital market imperfection that causes bills to be absent from the portfolio and money to act as the safe portfolio asset in addition to the transaction instrument. The Ando and Shell model describes an interior solution in a perfect capital market. The yield on bills represents the foregone interest of a wealth holder putting his wealth instead in money. The implicit assumption of this model is that the relevant tradeoff is independent of whether the wealth holder is a net issuer or holder of bills—the typical wealth holder is supposed to borrow (or invest) at the same rate as the U.S. Treasury.

We suggest, on the contrary, that there is a capital market imperfection, characterized as a spread between borrowing and lending rates available to a typical wealth holder. The existence of the spread is hardly in doubt; its causes probably lie in moral hazard and adverse selection issues beyond the scope of this paper. The result of the rate spread is that wealth holders face three regimes: debtor on the bill market, null position on the bill market, and creditor on the bill market. In the first and last cases, an interior solution obtains and the Ando–Shell analysis applies: the determination of transactions balances includes no role for risk. In the middle case, where the wealth holder has no net position on the bill market, since bills are not held or owed, the risk-return tradeoff on bonds enters the money-holding decision. The relevant tradeoff is safe money versus risky bonds: money fulfills both a transactions and a portfolio function, and measures of risk and return to holding financial assets enter explicitly in the money demand decision. We demonstrate that the very model that denies a risk-return tradeoff between money and risky assets in a perfect capital market requires it in an imperfect capital market. Hence, money demand specifications may need to include a risk-return tradeoff.

We start with the Ando and Shell (1975) model in a comparative static analysis. There are three assets: e (equity), s (savings deposits), and m (money) with yields $\tilde{\rho}$, r_s and r_m. $\tilde{\rho}$ is a random real rate of return, whereas r_s and r_m are certain nominal rates. The inflation rate, $\tilde{p}$, is random. e, s and m are expressed as average proportions of wealth: $e+s+m\equiv 1$.

Transaction costs T are the basis for money demand in the model, and vary negatively with m. The model has two periods with respective consumptions C_1 and C_2. Consumption is financed from wealth and investment yields. Portfolio selection is intended to maximize $EU(C_2)$ where E denotes the expectations operator and:

$$C_2 = W - C_1 + (W - \tfrac{1}{2}C_1)[e\tilde{\rho} + s(r_s - \tilde{p}) + m(r_m - \tilde{p})] - T(M, C_1). \tag{2}$$

C_1 (first period consumption) is taken as independent of the portfolio choice. W, initial wealth, is exogenous. Average money balance is a portion of average unconsumed wealth: $M = m(W - \frac{1}{2}C_1)$. Optimizing policy is described by choosing e and m to:

$$\text{Max}_{e,m}\, E_{\tilde{\rho},\tilde{p}}\, U(C_2). \tag{3}$$

The first order conditions, assuming an interior solution are:

$$\frac{\partial EU}{\partial e} = E\{(W - \tfrac{1}{2}C_1)[\tilde{\rho} - (r_s - \tilde{p})]U'(\cdot)\} = 0, \tag{4}$$

$$\frac{\partial EU}{\partial m} = E\{(W - \tfrac{1}{2}C_1)[-(r_s - \tilde{p}) + (r_m - \tilde{p} - T_m(\cdot))]U'(\cdot)\} = 0. \tag{5}$$

Simplifying, the second condition becomes:

$$r_s - r_m = -T_m(M, C_1). \tag{6}$$

Thus, Ando and Shell conclude that the money/savings-deposits margin is determined only by yield and transaction cost, while the equity/savings-deposit margin reflects risk and return: the risk-adjusted marginal utility of returns to s and e are equated.

We now complicate the model by restrictions to represent capital market imperfection:

$$\left.\begin{array}{ll} r_s = \bar{r}_s & \text{for } s < 0 \\ r_s = \underline{r}_s & \text{for } s > 0 \end{array}\right\} \quad \text{where } \bar{r}_s > \underline{r}_s.$$

That is, on the short-term money market, the individual can borrow only at a higher interest rate than that at which he can lend. This includes as a special case the impossibility of short-term borrowing: $\bar{r}_s = +\infty$. This formulation is in contrast with Ando and Shell where there are no non-negativity restrictions on s or m and there is a single borrowing and lending rate r_s.

Solving the optimization problem, there are three cases:

Case 1: $s > 0$; Case 2: $s = 0$; Case 3: $s < 0$.

For cases 1 and 3, the analysis of Ando and Shell remains, with $\underline{r}_s$, $\bar{r}_s$, substituted respectively for r_s. Case 2 is characterized by:

$$\bar{r}_s - r_m > -T_m > \underline{r}_s - r_m,$$

so the marginal returns to additional money holding neither reward indebtedness ($s < 0$, $r_s = \bar{r}_s$) nor savings deposits ($s > 0$, $r_s = \underline{r}_s$). Since $1 - e - m = 0$, $m = 1 - e$ and so:

$$U = U\{W - C_1 + (W - \tfrac{1}{2}C_1)[e\tilde{\rho} + (1-e)(r_m - \tilde{p})] - T[(W - \tfrac{1}{2}C_1)(1-e), C_1]\} \tag{7}$$

and hence the first-order condition is:

$$\frac{\partial EU}{\partial e} = E\{(W - \tfrac{1}{2}C_1)[(\tilde{\rho} - r_m + \tilde{p}) + T_m]U'(\cdot)\} = 0. \tag{8}$$

That is, the risk-adjusted marginal utility returns to m and e are equated. Hence, for case 2, the risk-return tradeoff between m and e (with the slight additional complication of T_m) is precisely parallel to that for s and e in cases 1 and 3. An important consequence of this formulation is that alternative asset yield and risk enter the money-demand function, so this formulation will differ from the Goldfeld benchmark when risk-adjusted yields on alternative assets differ from their average values.

Chou (1988) develops the comparative statics of case 2, for the special case of a quadratic utility function, $U(x) = ax + bx^2$. He finds the elasticity of m with respect to the variance of equity yields, $\sigma^2_{\tilde{\rho}}$, to be:

$$\mathscr{E}_{m,\sigma^2_{\tilde{\rho}}} = \left[\frac{e}{m}\sigma^2_{\tilde{\rho}} + \frac{1}{2}\left(\frac{e}{m} - 1\right)\sigma_{\tilde{\rho}\tilde{p}}\right] D^{-1} \tag{9}$$

where $\sigma_{\tilde{\rho}\tilde{p}}$ is the covariance of equity yields $\tilde{\rho}$ with inflation $\tilde{p}$, and:

$$D = (\mu_{\tilde{\rho}} - r_m - \mu_{\tilde{p}} + T_m)^2 + \sigma^2_{\tilde{\rho}} + \sigma^2_{\tilde{p}} + 2\sigma_{\tilde{\rho}\tilde{p}}$$
$$- \frac{T_{mm}}{b}\{a + b(W - C_1) + b(W - \tfrac{1}{2}C_1)[e\mu_{\tilde{\rho}} + m(r_m - \mu_{\tilde{p}})] - bT\}$$

and where $\mu_{\tilde{\rho}}$ and $\mu_{\tilde{p}}$ are the expected values of $\tilde{\rho}$ and $\tilde{p}$ respectively. In order to sign the expression for $\mathscr{E}$, if we disregard $\sigma_{\tilde{\rho}\tilde{p}}$ and the higher-order terms in D, we find that the elasticity of money demand with respect to equity risk is positive and proportional to equity risk. This is the formulation estimated below in the context of a dynamic equation.

Case 2 is best characterized by the yield inequality: $\bar{r}_s - r_m > -T_m > \underline{r}_s - r_m$. When the yield inequality holds, the expected return $\mu_{\tilde{\rho}}$ on e may exceed the return to savings, but be less than the cost of borrowing. The target level of bill holdings will be nil; equity rather than savings will be the preferred alternative for excess money balances. A wealth holder fulfilling the inequality may temporarily have a non-null position on the bill market—a transitory situation as asset balances are adjusted to target levels. When the yield inequality holds, the wealth holder will seek to pay off short-term debt and to run down short-term interest-bearing assets. This adjustment takes place through corresponding opposite adjustment in money-holding, m and in the equity position, e. The balance between m and e is then determined by a risk-return tradeoff. In the estimated money demand model below, long-term bonds are used to represent the risky asset e; their expected return is taken to be the yield to maturity and their risk is represented as a moving-average standard deviation of holding-period yield.

Inflation rates and interest rates both represent an opportunity cost of money-holding. Inflation rates, independently of and in addition to interest rates, will enter the money demand function if inflation and interest rates are imperfectly correlated, i.e. if the Fisher equation does not hold uniformly and without lag (see Figure 7 below). In the sample period 1960(3)–1988(3), the correlation of inflation rates and T-bill yields is 61% and so inflation rates account for $R^2 = 0{\cdot}37$ of the variation in T-bill yields. Hence, inflation is a separate argument in the money demand function, consistent with Goldfeld and Sichel (1990).

III. ECONOMETRIC MODELLING

The framework we adopt is discussed in Hendry and Richard (1982, 1983), Hendry (1987, 1989) and Spanos (1986). Empirical econometric models are viewed as derived representations of the economic mechanism, obtained by data reduction, using in-sample test statistics as model selection criteria. The main reduction steps in model derivation are:

[1] Data Transformation;
[2] Marginalization with respect to disaggregated and unwanted information;
[3] Sequential conditioning with respect to the history of the process to create a martingale difference error relative to the retained information;

[4] Linear approximation with a fixed lag length, selected to retain an innovation error;
[5] Conditioning on contemporaneous variables which are weakly exogenous for the parameters of interest (see Engle *et al.* (1983));
[6] Simplification to yield a parsimonious and interpretable data characterization.

Relative to the approach of simply fitting a theory-based model, [1]-[4] highlight the need for a complete menu of data-relevant determinants, an adequate dynamic specification (see Hendry *et al.* (1984)) and useful functional form transformations. Also, [5] precludes arbitrary exogeneity claims if the parameters of interest are to remain constant across regime shifts in the conditioning variables (see Engle and Hendry (1989)). However, [6] is *au choix*: general dynamic equations are susceptible to many possible simplifications even after the regressors have been expressed in nearly-orthogonal form. Simplification is usually necessary to interpret the model and to produce more robust specifications than unconstrained models. Further, parsimony allows more rigorous testing of the assumptions required to sustain the reductions.

The initial empirical model is formulated with sufficient generality to encompass previous findings, salient data features and theoretical knowledge, such that it would be surprising if a still more general model were necessary in order to adequately characterize the data. If the data base is insufficient to estimate the general model (e.g. the sample size is too small), a pre-simplified feasible general unrestricted model (denoted GUM) is estimated, to be expanded if necessary as data accrue. The data series are then transformed to create a near-orthogonal parameterization, selected to correspond to likely decision variables of the relevant economic agents, contingent on information they could have had available. The transformed model is simplified to eliminate redundant influences (which may be genuinely conditionally irrelevant, or may just lack variability in the given data set). Thus, the modelling strategy advocated designs congruent models by exhausting the information content of existing data so that such models cannot be dominated within-sample on that data.

Once a model is selected, an array of tests can be applied to check its congruency. The reduction steps delineate the information sets against which model validity can be checked:

(A_i) The past of the investigator's own data, leading to tests for homoskedastic innovation errors in the simplified conditional model ([3]).
(A_{ii}) The contemporaneous values of the conditioning variables, leading to tests of their weak exogeneity for the parameters of interest ([5]).
(A_{iii}) Future data, leading to tests for the constancy of the parameters of interest and hence their potential invariance (see Favero and Hendry (1989)).
(B) Theory information, leading to tests for theory consistency.
(C) Measurement information (e.g. inherent properties of the accounting system), and the associated concept of data admissibility.
(D) Complementary data used in rival models, again partitioned into the relative (i) past, (ii) present and (iii) future, leading to tests for historical encompassing (see Mizon (1984), Mizon and Richard (1986), and Hendry and Richard (1989)); exogeneity encompassing; and forecast encompassing (see Chong and Hendry (1986)) respectively.[2]

2. Goodness-of-fit is necessary but not sufficient for encompassing in a given model class.

The specific tests used to evaluate the various aspects of congruency comprised:

$\eta_1(M-N, T-K-M)$	F-test for N-th to M-th order residual autocorrelation in a model with K regressors and T observations: A_i (see Harvey (1981)).
$\eta_2(H, T-H-K)$	Chow F-test of parameter constancy over H forecasts; η_2^* is the analysis of covariance test ($H > K$): A_{iii} (see Chow (1960)).
$\eta_3(n, T-n-K)$	F-test of functional form mis-specification/heteroskedasticity for n variables: A_i (see White (1980)).
$\eta_4(m, T-m-K)$	F-test of the restricted model against the GUM for m restrictions: A_i.
$\xi_5(2)$	χ^2-test for Normality: A_i (see Jarque and Bera (1980)).
$\eta_6(r, T-2r-K)$	F-test for Autoregressive Conditional Heteroskedasticity (ARCH) of r-th order: A_i (see Engle (1982)).
$\eta_7(j, T-j-K)$	F-version of the RESET test for j powers: A_i (see Ramsey (1969)).
$\eta_8(x_t)(1, T-K-1)$	F-test on the significance of adding x_t as a regressor: A_i.
$\xi_9(H)/H$	χ^2-test for predictive failure over H forecasts, standardized by its degrees of freedom: A_{iii} (see Hendry (1979) and Kiviet (1987)).

Of these, η_1, η_3, η_4, η_6 and η_8 test for homoskedastic innovation errors; η_3 and η_7 for correct functional form specification; ξ_5 for normality; and η_2 and ξ_9 for constant parameters; also, η_4 and η_8 test for *a priori* restrictions to be imposed on the GUM. Such tests are part of the design strategy and despite being selection criteria are conducted in the Lagrange-Multiplier spirit (see Engle (1984)). In particular, while it is easy to remove (say) residual autocorrelation by re-designing the model, it seems difficult to produce parameter constancy just by tinkering with the specification (design) of a model so long as regime-shift dummies are not admissible. Even if this were feasible within sample, later evidence will accrue to discriminate good models from bad. In practice, the products of such an approach have tended to be parsimonious, robust and reasonably constant, as well as successfully accounting for the results of previous models (see Davidson *et al.* (1978) and Hendry (1979) *inter alia*).

Four aspects of our econometric approach deserve brief note, given the previous research on M1 demand in the U.S.A.: non-stationarity, autocorrelation, contemporaneous conditioning, and collinearity. We take these in turn.

Economic data are not well-characterized as being generated by a stationary stochastic process. First, the evolutionary nature of the levels of economic variables has prompted much recent research (see Hendry (1986), Engle and Granger (1987), Phillips (1991), Johansen (1988) and Hylleberg and Mizon (1989) *inter alia*). The model used below is consistent with the existence of a cointegrating relationship describing the long-run dependence of money demand on the arguments of the demand function. Denote the Vector Autoregression (VAR) of order p in the N variables $\{\boldsymbol{x}_t, t=1, \ldots, T\}$ under analysis by:

$$\Delta \boldsymbol{x}_t = \boldsymbol{\phi} + \textstyle\sum_{i=1}^{p-1} \Pi_i \Delta \boldsymbol{x}_{t-i} + \Pi \boldsymbol{x}_{t-p} + \boldsymbol{e}_t \quad \text{where } \boldsymbol{e}_t \sim IN(\mathbf{0}, \boldsymbol{\Omega}).$$

The $\boldsymbol{x}_t$ are assumed to be integrated of order one, $I(1)$, but the number of cointegrating vectors ν, and hence the number of unit roots $N-\nu$, is unknown. The distributions of statistics are non-standard in such a setting and require special critical values. Tests for ν can be based on the approach proposed by Johansen (1988), which is equivalent to testing whether $\Pi = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ are $N \times \nu$. Since the likelihood function depends on the normal distribution, it can be concentrated with respect to $\boldsymbol{\Omega}$, $\boldsymbol{\phi}$ and $(\Pi_1, \ldots, \Pi_{p-1})$, the last by obtaining residuals $\boldsymbol{U}_{0t}$ and $\boldsymbol{U}_{pt}$ from regressing $\Delta\boldsymbol{x}_t$ and $\boldsymbol{x}_{t-p}$ respectively on the $\{\Delta\boldsymbol{x}_{t-i}\}$. Denote the second moment matrices from these residuals by $\boldsymbol{S}_{00}$, $\boldsymbol{S}_{0p}$, $\boldsymbol{S}_{p0}$ and $\boldsymbol{S}_{pp}$ where $\boldsymbol{S}_{ij} = T^{-1}\sum_{t=1}^{T} \boldsymbol{U}_{it}\boldsymbol{U}'_{jt}$ for $i, j = 0$ and p. Then ν is determined by the largest eigenvalues $\lambda_1 \geqq \cdots \geqq \lambda_\nu \geqq \cdots \geqq \lambda_N \geqq 0$ of:

$$|\lambda \boldsymbol{S}_{pp} - \boldsymbol{S}_{p0}\boldsymbol{S}_{00}^{-1}\boldsymbol{S}_{0p}| = 0,$$

and $\hat{\boldsymbol{\beta}}$ by the corresponding eigenvectors. Since the maximized likelihood is given by:

$$L(\hat{\boldsymbol{\beta}}) = C - \tfrac{1}{2}T\sum_{i=1}^{\nu} \ln(1-\lambda_i),$$

tests of the hypothesis that there are $0 \leqq \nu < N$ cointegrating vectors can be based on:

$$\xi_{10}(\nu) = -T\sum_{i=\nu+1}^{N} \ln(1-\lambda_i),$$

(twice the log of the likelihood ratio for restricting Π) with ν being selected via the first significant statistic $\xi_{10}(\nu)$. Alternatively, sequential tests of significance of the largest $\{\lambda_\nu\}$ can be based on $\xi_{11}(\nu) = -T\ln(1-\lambda_\nu)$. Under the null hypothesis that the eigenvalues are zero, both $\xi_{10}(\nu)$ and $\xi_{11}(\nu)$ have distributions which are functionals of Brownian motion, but critical values for these tests have been tabulated by Johansen and Juselius (1990) and Osterwald-Lenum (1990) *inter alia*. The cointegrating combinations are given by $\hat{\boldsymbol{w}}_t = \hat{\boldsymbol{\beta}}'\boldsymbol{x}_t$ and these are the estimated error-correction mechanisms (ECMs). Further, $\hat{\boldsymbol{\alpha}}$ reveals the importance of each cointegrating combination in each equation. If a given ECM enters more than one equation, the parameters are cross-linked between such equations, violating weak exogeneity and entailing joint estimation for efficient estimation (see Phillips and Loretan (1991) and Hendry and Mizon (1990)). We exclude the possibility that $\boldsymbol{\phi}$ only enters the cointegrating vectors, since GNP has grown substantially over the sample period.

Second, although all the variables in our model are either differenced or cointegrated, stationarity is not a reasonable basis for modelling, given the predictive failure symptomatic of previous M1 relationships in the U.S.A. (for a related analysis applied to the U.K., see Hendry and Ericsson (1991)). Regime shifts, financial innovation and structural change all offer potential explanations for previous model failures. Thus, our primary focus is to find constant parameter representations, which incorporate past changes and are formulated to remain constant despite changes elsewhere in the economic system (see Hendry (1979, 1985)).

Next, the presence of autocorrelated residuals is often "treated" by assuming an underlying autoregressive error process and "removing" it via one of the many packaged devices (such as that attributed to Cochrane and Orcutt (1949)). This is not a valid approach to dynamic specification in general. As shown by Sargan (1980) and exposited by Hendry and Mizon (1978), autoregressive errors entail a range of testable restrictions on the general dynamic model (called common factor restrictions) and these should be tested first, prior to fitting the error process. In many cases, common factor restrictions are rejected when tested, usually because the residual autocorrelation is a symptom of model mis-specification (often incorrect dynamics), not of autoregressive errors. Conversely, when common factors are a valid representation, the data should be consistent with the restrictions (for examples, see Mizon and Hendry (1980) and McAleer *et al.* (1985)). This is one important reason for commencing from a general dynamic model.

The simultaneous-equations paradigm is so dominant in textbooks that many economists seem skeptical of models with contemporaneous conditioning variables, i.e. they doubt the validity of weak exogeneity assertions. A proper test of valid conditioning would involve formulating a congruent marginal model and then testing for the dependence of the parameters of interest on the parameters of the marginal model. Since one point of formulating conditional models rather than jointly modelling every variable is because of the difficulty of doing the latter, direct testing of weak exogeneity is rarely practical. Instead, necessary conditions can be tested, including the absence of the long-run money demand cointegrating vector from the other marginal models (see Hendry and Mizon (1990)). Further, since weak exogeneity is necessary for super exogeneity (see Engle and Hendry (1989)), and since the behaviour of some of the explanatory variables must have altered over the sample when predictive failure has in fact occurred, parameter constancy tests indirectly test weak exogeneity. If the conditional model does have constant parameters despite other models having failed, then the evidence is consistent with valid conditioning. When a set of valid instrumental variables exists, of which no current or lagged values enter the conditional model, then lags of those instruments can be used to test the overall validity of the model formulation and indirectly that of the conditioning assumptions. Below, the legitimacy of conditioning is tested via cointegration and sub-sample constancy tests.

Moreover, when the resulting conditional demand model is constant, but it is known that the supply function shifted during the sample period (e.g. under the New Operating Procedures for money supply), then the "classical" identification problem does not arise since all linear combinations involving the shifting equations are automatically excluded. Thus, a constant relationship with appropriately interpretable coefficients could only be a demand function (compare Cooley and LeRoy (1981)).

Finally, collinearity is not a property of a linear model, but of a parameterization of that model. The important issues thus concern the choice of the parameters of interest and the information content of the data, and not the correlations between some formulation of the regressors. For interpretability, parameters of interest should correspond to relatively orthogonal variables and for parsimony, those functions of the data which lack variability should be deleted. Since any linear model can be orthogonalized, collinearity at the level of the general unrestricted model is not a problem, subject to numerical accuracy *caveats* which sensible scaling and good programming can mitigate.

The economic theory model delineates the *class* of econometric models to be investigated, namely, error-correction formulations with well-defined equilibrium properties entailing that a cointegrating vector exists. The log-linear functional form was selected to ensure that the model was equivalent under a range of transformations of the dependent variable (including using (log) nominal or real money or velocity or changes in any of these), with potentially homoskedastic residuals (as a percentage of real M1), which seemed likely to capture the constant behavioural propensities of economic agents in an evolving world. Since long lag lengths on all variables need not entail slow lag reactions, the longest lag was set at six initially (on money, prices and incomes, but one or two on interest rates), producing 40 parameters, which was acceptable given a sample of around 100 observations.

The original specification was developed in Baba, Hendry and Starr (1985) (denoted BHS below), who transformed their general model to a near-orthogonal parameterization (fitted using the sample up to 1984(2)) and sequentially deleted redundant regressors. An F-test against the general model was calculated to protect against imposing invalid restrictions. So far as possible, individual regressors were retained only if they were "significant" at around the 0·1% level ($|t| \geqq 3{\cdot}46$ for more than 50 degrees of freedom) as this allowed at least 50 tests on the model while retaining the overall significance level at less than 5%. However, if deleting a sub-set of variables produced large values for diagnostic checks, the resulting variables were retained if their absolute t values exceeded 2·5.

Because a number of features of our model were relatively novel, there was little previous empirical evidence to narrow down the initial range of reasonable hypotheses. Consequently, we arbitrarily imposed certain formulations, leaving improvements on our initial specifications to later investigators should our approach prove useful. The simplified model was then evaluated against most of the information sets described above. Since that study, new data has accumulated, including eleven observations outside the longest sample previously used even for testing, and occurring over the M1 explosion episode noted above.[3]

IV. FORMULATING THE MONEY DEMAND FUNCTION

The error-correction model posits a long-run stable demand function for M1, deviations from which induce adjustments to re-establish equilibrium. Given the theoretical discussion in Section II above, the postulated long-run equation has the form:

$$\frac{M}{P} = \alpha_0 Y^{\alpha_1} \exp\left(\alpha_2 S + \alpha_3 R_1 + \alpha_4 \dot{p} + \alpha_5 R_{nsa} + \alpha_6 V + \alpha_7 R_{ma}\right) \tag{10}$$

where $M = \text{M1}$, $P =$ implicit deflator of GNP, $Y =$ real GNP, $S =$ Bond/T-bill yield spread, $R_1 =$ one month T-bill rate, $\dot{p} =$ the inflation rate, $R_{nsa} =$ learning-adjusted yield in M1 (to account for the introduction of interest-bearing NOW and SuperNOW accounts), $V =$ moving standard deviation of holding period yields on long-term bonds (measuring the risk to the holding of long-term debt) and $R_{ma} =$ learning-adjusted maximum yield in M2 (accounting for the introduction of new assets and institutional change). Lower case letters denote the logs of the corresponding capitals. Figure 1 reports the time series of $(m-p) = \ln M/P$ over the entire sample; note the dramatic change in the behaviour of the series at the end of the sample. Figure 2 records the time series of inverse velocity $(m-p-y)$ which reveals that the ratio fell steadily during 1960–1984, before rising at the end of the sample.

The specification in (10) is consistent with, and generally nests, previous long-run solutions as in (1) above. Although the distinctive economic variables in the present demand function all have antecedents in the literature,[4] three variables are sufficiently novel to merit separate discussion, namely V, R_{ma} and R_{nsa}.

First, the expected risk and return to long-term bonds must be modelled. Assuming expected capital gains from anticipated movements in long-term interest rates to be nil, the expected return is simply the yield. The expected risk is represented as a slowly evolving measure, denoted V, rather than a highly variable measure (such as ARCH or the innovation variance of the holding-period yield H_t), which might be appropriate in a speculative market (see Pagan and Ullah (1985)). Thus, we summarize risk by a moving standard deviation of H_t (at monthly rates). Let:

$$\bar{H}_t = \left(\textstyle\sum_{i=1}^{12} H_{t-i}\right)/12. \tag{11}$$

3. In fact, the data series on prices and income were substantially revised to a 1982 basis, which induced some changes in the empirical dynamics of M1 demand, and revealed a simpler specification for interest rates than that reported in BHS. The measure of volatility was amended to be an average over two years of the annual standard deviation of holding period yields at a quarterly rate. The M1 data have been revised recently also, but we retained the earlier data for comparability. Similar results were obtained on the newer figures, except for two outliers not present on the original data set.

4. See *inter alia* Evans (1984), Koskela and Viren (1984), Rose (1985), Starr (1983), Tatom (1984), and Turk (1984). The actual measures adopted here as corresponding to the theoretical variables undoubtedly remain open to improvement, but as we show below, account for most previous failures.

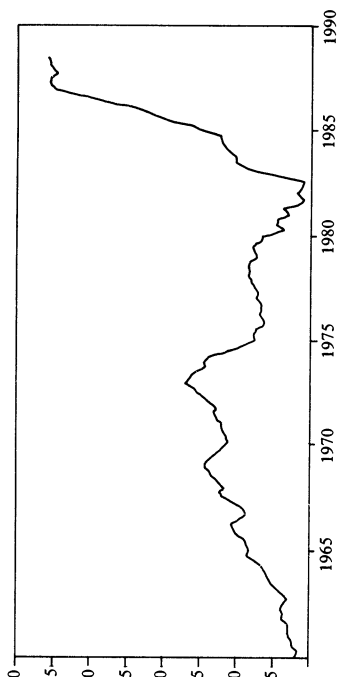

FIGURE 1
Time series of log real money $(m-p)$

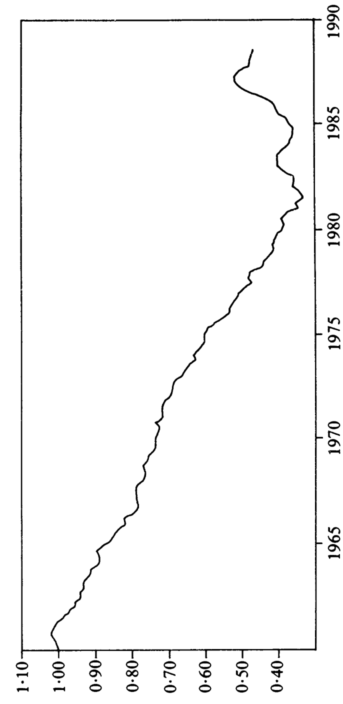

FIGURE 2
Time series of log inverse velocity $(m-p-y)$

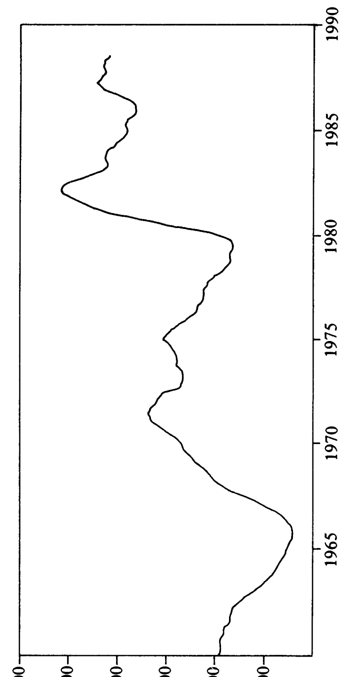

FIGURE 3
Time series of volatility V

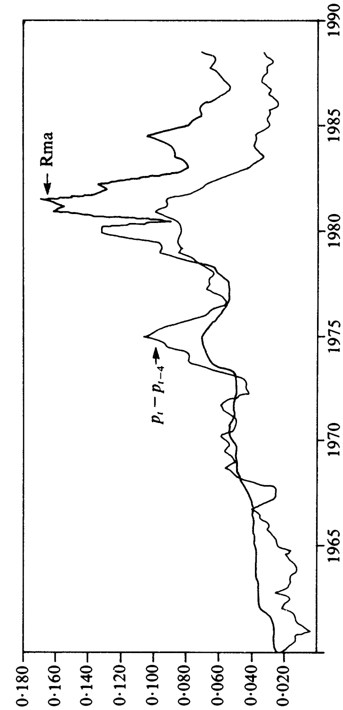

FIGURE 4
Time series of interest rates and annual inflation

Representation of bond holding-period risk is first by an annual standard deviation of H_t. In month t, define:

$$V_{m_t} \equiv \left(\sum_{j=1}^{12} (H_{t-j} - \bar{H}_{t-j})^2/12\right)^{1/2}. \tag{12}$$

Then V_m is converted to a quarterly average V_q, from which our measure V is calculated as the moving average:

$$V_t \equiv \sum_{i=0}^{8} V_{qt-i}/9.$$

A graph of the behaviour of $\{V_t\}$ over time is provided in Figure 3, and a listing of V_{mt}, V_{qt} and V_t is available from the authors on request.[5] Note that V_t has a local trough in 1973(4) and a local peak in 1981(4). These dates are immediately prior to periods of structural breakdown in most other M1 models, which are generally characterized by over-prediction of demand in 1974–1976 and under-prediction in 1982–1983. We regard this as more than simple coincidence, and V_t is important to the model's ability to forecast these periods of known difficulty and to predict the major upswing in M1 in 1985–1987. Since the growth rate of nominal money exhibits considerably varying volatility, a check on the adequacy of V_t is provided by testing for ARCH effects in the error term of our finally selected model (see Engle (1982) and Engle and Hendry (1989)).[6] One interpretation of the role of risk and return to long-term bonds in money demand is that the relevant long-term bond yield is a "certainty equivalent" yield on long-term bonds represented as the bond yield minus the premium in the yield attributable to asset price risk. This interpretation is developed in the model of equations (21) and (22) below.

Next, much financial innovation during the 1970's and 1980's has taken the form of creating new financial instruments which are close substitutes for other monetary instruments but which typically carry higher yields (see e.g. Simpson and Parkinson (1984)). A period of time is generally required for wealth holders to learn about, adapt to and trust a new instrument. Thus, the effective marginal opportunity cost of holding wealth in non-interest bearing M1 (as opposed to interest-bearing M2) can be represented by the differential from the highest prevailing yield in non-transactions M2, with the yield on newer instruments being weighted by a learning curve. A period of 20 quarters is assumed sufficient for full effectiveness of the new instrument. The learning curve reflects both pure information accrual as to the existence of the asset (an "epidemic" ogive response curve as in Griliches (1957)) and the economic motivation for learning to use it (a function of its excess rate of return over similar existing assets). Again, we have little previous research to build on, so given the number of features to be data modelled, we selected a plausible if *ad hoc* specification. Let t_0 be the introduction date of the asset, then the learning effect is:

$$w(t) = \begin{cases} 0, & t < t_0 \\ [1+\exp(7{\cdot}0 - 0{\cdot}8(t - t_0 + 1))]^{-1}, & t_0 \leqq t < 20 + t_0 . \\ 1, & t \geqq 20 + t_0 \end{cases} \tag{13}$$

5. The lag dependence of V_t on H_t was not subject to experimentation except for allowing up to 2 lags in the GUM, and only V_t itself was retained in the final model. For comparison, we computed a 36-month moving standard deviation (suggested by Neil Ericsson) and obtained similar but somewhat poorer results. V_t relatively downweights very recent and long past values of H_t compared to a pure moving standard deviation.

6. Since the residuals from a time-series model for Δm_t have a first-order ARCH coefficient of 0·44 ($t = 4{\cdot}9$) whereas our model has no ARCH effects in its residuals, our treatment of risk and dynamics is effective in accounting for that data feature among others.

The function in (13) is designed to be negligible prior to the asset being introduced, and to be unity at t_0+20.

The raw yields on non-transactions M2 assets considered here are:

R_p = passbook rate
R_{cd} = commercial bank small certificate of deposit (cd) rate
R_{mf} = money market mutual fund rate.

The dates for the weights in (13) are, therefore: $t_0 = 1965(4)$ for small bank *cds*, w_{cd}; and $t_0 = 1974(3)$ for money market mutual funds, w_m. From these weights, a family of adjusted yields representing learning and accommodation by wealth holders to new M2 assets is generated using (13):

$$R_{cda} = R_p + w_{cd}(R_{cd} - R_p)$$

$$R_{mfa} = R_{cda} + w_m(R_{mf} - R_{cda})$$

from which

$$R_{ma} = \max\,[R_p, R_{cda}, R_{mfa}].$$

R_{ma} is the representative learning-adjusted yield on non-transactions M2. This approach generates a continuous time series $\{R_{ma_t}\}$ for the alternative asset yield (essential for econometric use) and at least reflects the main ingredients of choice in an evolving financial environment. Figure 4 shows R_{ma_t} and annual inflation $(p_t - p_{t-4})$.

Next, we applied the same treatment to the introduction of interest-bearing M1 assets:

R_n = NOW account rate, taken to be nil prior to national introduction in 1981(1).
R_{sn} = SuperNOW rate, taken to be nil prior to national introduction in 1983(1).

For M1 assets, the dates in (13) are given by: $t_0 = 1981(1)$ for NOW accounts, w_n; and $t_0 = 1983(1)$ for SuperNOW accounts, w_{sn}. Thus, we create:

$$R_{na} = w_n R_n$$

and

$$R_{sna} = w_{sn} R_{sn}$$

from which

$$R_{nsa} = \tfrac{1}{2}[R_{na} + R_{sna}].$$

Then R_{nsa} is the adjusted other-checkables rate in M1.[7]

The variables and data sources are fully described in Appendix A but in summary are:

m_t = ln M1, seasonally adjusted.
p_t = ln GNP deflator, seasonally adjusted.
y_t = ln real GNP, seasonally adjusted.

7. Both the data on the flows into NOW and SuperNOW accounts and the early estimates of separate effects for R_{na} and R_{sna} (see e.g. Table 1) suggested that these assets appealed to different clienteles; hence the use of an average.

$Ay_t = \frac{1}{2}(y_t + y_{t-1})$ (an average suggested by Rose (1985)).
R_t = 20-year Treasury bond yield to maturity.
R_{1_t} = 1-month Treasury bill coupon equivalent yield.
$AR_{1_t} = \frac{1}{2}(R_{1_t} + R_{1_{t-1}})$.
$S_t = R_t - R_{1_t}$.
$AS_t = \frac{1}{2}(S_t + S_{t-1})$.
R_{ma_t} = learning-adjusted maximum yield on instruments in M2 (see text).
$R_{nsa_t} = \frac{1}{2}(R_{na_t} + R_{sna_t})$: learning-adjusted other-checkables rate in M1 (see text).
V_t = volatility measure based on long-bond holding-period yields (see text).
$SV_t = \max(0, S_t) \cdot V_t$.
D_t = credit control dummy, -1, $+1$ in 1980(2)/(3), 0 otherwise (see Gordon (1984)).
$\Delta_i x_t = (x_t - x_{t-i})/i$ for any variable x_t.
$\Delta^2 x_t = \Delta x_t - \Delta x_{t-1}$.
$\Delta \hat{p}_t = \Delta p_t + \Delta^2 p_t$ (predictor of inflation: see Campos and Ericsson (1988)).

All interest rates are expressed as decimal fractions, at annual rates, so V_t is scaled by 0·01.

V. AN EMPIRICAL MONEY DEMAND MODEL

To establish the orders of integration of the processes in the model and check the cointegration entailed by (10), the six main variables $(m_t - p_t, \Delta p_t, y_t, V_t, S_t, R^*_{1_t})$ were analysed using the procedure proposed by Johansen (1988), where $R^*_{1_t} = R_{1_t} - \frac{1}{2}R_{nsa_t}$. A VAR with 5 lags on each variable (plus intercepts) yielded the following eigenvalues $\{\lambda_i\}$ and eigenvectors $\hat{\boldsymbol{\beta}}$ of the long-run matrix $\Pi = \boldsymbol{\alpha}\boldsymbol{\beta}'$ for T = 1960(3)–1988(3):

Eigenvalues of Π

ν	λ_ν	$\xi_{11}(\nu)$	5% CV	$\xi_{10}(\nu)$	5% CV[8]
6	0·027	3·12	3·76	3·12	3·76
5	0·056	6·51	14·07	9·63	15·41
4	0·106	12·60	20·97	22·23	29·68
3	0·137	16·60	27·07	38·83	47·21
2	0·186	23·27	33·46	62·10	68·52
1	0·396	56·97	39·37	119·06	94·16

The null of no cointegrating vectors can be rejected at the 5% level in favour of one according to both tests $\xi_{10}(\nu) = -T\sum_{i=\nu+1}^{N} \ln(1-\lambda_i)$ and $\xi_{11}(\nu) = -T\ln(1-\lambda_\nu)$, but one cannot be rejected in favour of two. The first cointegration vector $\hat{\boldsymbol{\beta}}$ normalized on $(m-p)_t$ is:

$$(m-p)_t = -5{\cdot}51\Delta p_t + 0{\cdot}51 y_t - 6{\cdot}64 S_t - 3{\cdot}96 R^*_{1_t} + 3{\cdot}72 V_t. \tag{16}$$

8. We are grateful to Michael Osterwald-Lenum for permission to quote the critical values of the Johansen statistics. When interpreting this analysis, it must be remembered that (a) the coefficient of R_{nsa} was imposed from (18) at $\frac{1}{2}$ due to being non-zero for only a small part of the sample; (b) the VAR coefficients on some of the marginal processes were not constant; (c) we imposed the restrictions (from re-interpreting BHS) that R_{ma} and SV only entered the analysis in differenced form. All of these *caveats* entail that the VAR analysis is suggestive rather than definitive, hence we neglect the possible dependence of the spread equation on the cointegrating money demand vector.

This is consistent with the postulated long-run solution in (10) and the earlier results in BHS, and supports a coefficient of one half on income in the ECM. The feedback coefficients were:

$\boldsymbol{\alpha}'$-vector

Variable	$(m-p)_t$	Δp_t	y_t	S_t	$R^*_{1_t}$	V_t
$(m-p)_t$	−0·24	0·05	0·09	−0·14	0·08	−0·00

The first element in $\boldsymbol{\alpha}$ implies a feedback coefficient of −0·24 which is much larger than in BHS (but similar to that found on the revised data below). However, some of the other elements in the first column are large and suggest that the first cointegrating vector enters at least the spread equation as well, which would violate weak exogeneity.

Turning to the single-equation approach, based on the experience in BHS and earlier evidence of a single cointegrating vector, the GUM allowed up to six lags on (m_t, p_t, y_t), two lags on $(SV_t, R_{na_t}, R_{sna_t})$, and one lag on $(V_t, R_t, R_{1_t}, R_{ma_t})$, as well as including D_t (unlagged) and an intercept. Table I reports the GUM estimates together with F-tests on the overall relevance of each variable (with n_i and $T-K$ degrees of freedom for n_i lags in the given polynomial and $K=39$ regressors), and t-tests of the hypothesis that the sum of each variable's lag coefficients is zero (corresponding to a unit root in the associated lag polynomial: see Hendry (1989)). Both null hypotheses are rejected at the 5% level for $(m_t, p_t, y_t, V_t, R_t)$ (using a critical value of 4·8, based on 6 regressors in

TABLE I

General unrestricted model

Lag	0	1	2	3	4	5	6	Σ	F/t
m_t	−1·000	1·039	−0·381	−0·000	−0·133	0·414	−0·166	−0·228	92·90
SE	—	0·105	0·135	0·119	0·106	0·104	0·072	0·036	−6·39
y_t	0·194	−0·010	−0·044	−0·087	0·110	−0·055	0·023	0·131	8·64
SE	0·060	0·077	0·072	0·069	0·069	0·069	0·053	0·021	6·35
p_t	0·326	−0·221	−0·117	0·042	0·150	−0·134	0·172	0·218	8·01
SE	0·122	0·168	0·161	0·164	0·170	0·179	0·123	0·031	7·04
V_t	0·484	0·271	—	—	—	—	—	0·755	20·73
SE	0·436	0·437	—	—	—	—	—	0·117	6·43
R_t	−0·623	−0·607	—	—	—	—	—	−1·230	12·00
SE	0·239	0·293	—	—	—	—	—	0·259	−4·75
R_{1_t}	0·195	0·119	—	—	—	—	—	0·315	1·28
SE	0·191	0·217	—	—	—	—	—	0·205	1·53
R_{ma_t}	−0·226	0·297	—	—	—	—	—	0·071	5·10
SE	0·100	0·094	—	—	—	—	—	0·084	0·84
R_{na_t}	1·018	−1·465	0·746	—	—	—	—	0·298	2·88
SE	0·528	0·817	0·457	—	—	—	—	0·108	2·77
R_{sna_t}	−0·498	1·173	−0·500	—	—	—	—	0·175	1·72
SE	0·774	1·367	0·758	—	—	—	—	0·161	1·08
SV_t	1·28	9·45	−12.70	—	—	—	—	−1·97	10·10
SE	5·39	6·34	2·60	—	—	—	—	7·22	−0·27
D_t	0·013	—	—	—	—	—	—	0·013	13·54
SE	0·004	—	—	—	—	—	—	0·004	3·68
1	0·346	—	—	—	—	—	—	0·346	22·43
SE	0·073	—	—	—	—	—	—	0·073	4·74

$R^2=0·99996$, $\hat{\sigma}=0·398\%$, $F(38, 74)=49{,}422$, $DW=2·20$, $SC=-9·84$

$\eta_3(37, 36)=0·62$, $\xi_5(2)=0·22$, $\eta_6(4, 66)=0·59$, $T=1960(3)\text{–}1988(3)$

DW is the Durbin–Watson (1950, 1951) statistic and *SC* is the Schwarz criterion (see e.g. Hendry (1989)).

MacKinnon (1990), for the unit root t-test, which has a Dickey-Fuller (1979, 1981) distribution). This outcome matches the results of the preliminary cointegration analysis.

The solved long-run equation deriving from the Table I model is:

$$\underset{(0\cdot13)}{m = 1\cdot52} + \underset{(0\cdot045)}{0\cdot955p} + \underset{(0\cdot060)}{0\cdot574y} + \underset{(0\cdot52)}{3\cdot31V} - \underset{(0\cdot97)}{5\cdot39R} + \underset{(0\cdot83)}{1\cdot38R_1} + \underset{(0\cdot37)}{0\cdot310R_{ma}}$$
$$+ \underset{(0\cdot52)}{1\cdot31R_{na}} + \underset{(0\cdot627)}{0\cdot765R_{sna}} - \underset{(32\cdot2)}{8\cdot63SV} + \underset{(0\cdot020)}{0\cdot059D} \qquad (17)$$

This equation extends, but otherwise closely matches, the solution from the VAR.

Obvious restrictions on the GUM, such as price homogeneity, an income elasticity of 0·5, and the long-run irrelevance of R_{ma} and SV were tested (and accepted). The equation was then transformed as discussed in Section III. Short moving averages were formed in the light of the sample evidence, taking account of earlier findings. The following model for M1 demand was developed from the Table I GUM and constitutes the baseline equation discussed in the remainder of the paper. Standard errors are in parentheses, estimation was by OLS, the sample size was 113 observations, and Mean and SD respectively denote the dependent variable's unconditional mean and standard deviation.

$$\Delta(m-p)_t = \underset{(0\cdot020)}{0\cdot352} - \underset{(0\cdot104)}{1\cdot409AS_t} - \underset{(0\cdot063)}{0\cdot973AR_{1_t}} - \underset{(0\cdot049)}{0\cdot255\Delta R_{ma_t}} + \underset{(0\cdot070)}{0\cdot395\Delta Ay_t}$$
$$- \underset{(0\cdot046)}{0\cdot330\Delta\hat{p}_t} - \underset{(0\cdot132)}{1\cdot097\Delta_4 p_{t-1}} + \underset{(0\cdot055)}{0\cdot435R_{nsa_t}} + \underset{(0\cdot079)}{0\cdot859V_t}$$
$$+ \underset{(1\cdot49)}{11\cdot68\Delta SV_{t-1}} - \underset{(0\cdot015)}{0\cdot249(m-p-\tfrac{1}{2}y)_{t-2}} - \underset{(0\cdot097)}{0\cdot334\Delta_4(m-p)_{t-1}}$$
$$- \underset{(0\cdot039)}{0\cdot156\Delta^2(m-p)_{t-4}} + \underset{(0\cdot003)}{0\cdot013D_t} \qquad (18)$$

$R^2 = 0\cdot894$, $\hat{\sigma} = 0\cdot385\%$, $F(13, 99) = 64\cdot50$, $DW = 1\cdot89$, $SC = -10\cdot67$

Mean $= 0\cdot3\%$, $SD = 1\cdot11\%$, $T = 1960(3)\text{–}1988(3)$

$\eta_1(5, 94) = 0\cdot40$ $\eta_2(11, 88) = 0\cdot33$ $\eta_2(16, 83) = 1\cdot20$

$\eta_3(26, 72) = 0\cdot47$ $\eta_4(25, 74) = 0\cdot74$, $\xi_5(2) = 0\cdot45$

$\eta_6(4, 91) = 0\cdot44$ $\eta_7(1, 98) = 0\cdot13$ $\eta_8(y_t)(1, 98) = 0\cdot17$

$\eta_8(p_t)(1, 98) = 0\cdot48$ $\eta_8(R_{ma_t})(1, 98) = 0\cdot03$.

The interpretation of (18) is easiest by first examining the properties of the equilibrium solution and then analysing the dynamic adjustment. The derived equilibrium solution of (18) is close to that in (16) and (17):[9]

$$(m-p) = 1\cdot41 + 0\cdot5y + 3\cdot45V - 5\cdot65S - 3\cdot90R_1 - 1\cdot43\dot{p}_a + 1\cdot74R_{nsa} \qquad (19)$$

9. To estimate the standard errors of the derived coefficients in (19) and the response mean and median lags, the dynamics on m-p and y were derestricted, maintaining other restrictions, inducing the solution:

$$m-p = \underset{(0\cdot014)}{1\cdot40} + \underset{(0\cdot031)}{0\cdot53y} + \underset{(0\cdot31)}{3\cdot33V} - \underset{(0\cdot28)}{5\cdot85S} - \underset{(0\cdot14)}{3\cdot95R_1} - \underset{(0\cdot27)}{1\cdot62\dot{p}_a} + \underset{(0\cdot33)}{1\cdot62R_{nsa}} - \underset{(0\cdot29)}{1\cdot37\Delta R_{ma}}$$
$$+ \underset{(7\cdot9)}{51\cdot9\Delta SV} + \underset{(0\cdot016)}{0\cdot031D}.$$

Such calculations probably overestimate the uncertainty in (19).

where $\dot{p}_a$ is expressed as an annual rate for comparability with the annual interest rates. Assuming that all of the variables are indeed $I(1)$, then (18) both estimates the extended cointegration vector, and explains its existence as the outcome of error-correction behaviour by agents seeking to control their transactions balances in relation to income, inflation and a vector of competing and complementary interest rates, allowing for risk.

From (19) we can derive the following long-run response elasticities of the demand for money relative to the competing yield interest rate; the own yield interest rate; inflation; bond interest rate volatility; (long-short) yield spread; and income respectively:

$$\mathscr{E}_{(M/P),R_1} = -3{\cdot}9R_1 \qquad \mathscr{E}_{(M/P),R_{nsa}} = 1{\cdot}7R_{nsa} \qquad \mathscr{E}_{(M/P),\dot{p}_a} = -1{\cdot}4\dot{p}_a$$

$$\mathscr{E}_{(M/P),V} = 3{\cdot}4V \qquad \mathscr{E}_{(M/P),S} = -5{\cdot}7S \qquad \mathscr{E}_{(M/P),y} = 0{\cdot}5 \tag{20}$$

The T-bill rate elasticity ranges from −39% when the yield is 10%, to −20% when the yield is 5%; the own yield elasticity is approximately 12% when the maximal own-yield on M1 instruments is 7% (this seems small in view of the associated liquidity characteristics; the arbitrarily posited timing of the learning adjustment may be unreliable); the inflation elasticity is roughly −14% at 10% inflation;[10] the volatility elasticity is approximately 17% at V's peak value of 0·051, or 9% at its mean of 0·026; the spread elasticity is near zero at S's minimal value of −0·005 and almost −24% at S's maximum value of 0·04 (which is large but has an interpretable sign, while entailing a positive coefficient on R_1 in terms of the level of the interest rate: compare (17)); finally the long-run income elasticity of demand is 0·5, consistent with the simplest version of the Baumol (1952)–Tobin (1956) theory; although the coefficient of one half is imposed in the error correction, it is data-coherent, being 0·51 if derestricted in (18) (whereas unity is data-rejected). Thus, the equation supports the earlier argument that the additional and reformulated variables are potentially important.

Turning next to the interpretation of the dynamic equation (18), it must be stressed that it is a conditional model, both formulated and estimated as a contingent function of the information available to wealth holders in adjusting their portfolios and transactions demands. It can be interpreted either in terms of real or of nominal money demand, but because of the error-correction formulation, long-run demand fully adjusts to price level changes, as exhibited in (19). The results are consistent with the class of target-bounds models described by e.g. Miller and Orr (1966), Akerlof (1979), Milbourne (1983), and Smith (1986), with short-run nominal adjustment being within bounds set by long-run real magnitudes.

The dynamic adjustment is complicated, with responses to an annual average of the lagged dependent variable ($\Delta_4(m-p)_{t-1} \equiv \frac{1}{4}\sum_{i=1}^{4} \Delta(m-p)_{t-i}$), its second difference at $(t-4)$ (which may compensate for separate seasonal adjustment of the other regressors), and the primary level feedback at $t-2$. Nevertheless, both the mean and median lag responses to income are rapid, being respectively about 3 and just over 1 quarter.[11] These

10. The coefficient of annual inflation in (19) is in the mid-range of Cagan's (1956) values when expressed in equivalent time-units. Cagan's treatment of seven hyper-inflations includes both an exponential response to (expected) inflation rates and an exponentially distributed lag for converting actual to expected inflation rates. As a monthly rate, Cagan's exponential response coefficient is estimated from 2·3 to 8·7, depending on the country, with 90% confidence intervals ranging from 1·7–3·9 to 6·4–42·2. Converting the present coefficient (1·43 at annual rate) to monthly gives 17·2. Hence the present estimate, at inflation rates lower by several orders of magnitude, lies in the same range. Here, the expectations mechanism uses the data-based predictor $\Delta p_t + \Delta^2 p_t$, but the major role of inflation in the equation is played by $\Delta_4 p_{t-1}$.

11. Calculated as described in footnote 9. The lag weight distributions entailed by the dynamics of $(m-p)$ are oscillatory, and induce small negative weights at long lags in some of the reaction functions: this *caveat* affects the interpretation of summary statistics of lag distributions.

are far shorter than recorded in models such as (1), where a long mean-lag is often an artifact of the restricted dynamic specification due to assuming a common exponential decay rate.

The competing interest rate effects comprise a reaction of about -1 to a short moving average of the T-bill rate, of $-1{\cdot}15$ to the moving average long-short spread and of $-\frac{1}{4}$ to the *change* in the maximal learning-adjusted interest rate on non-transactions M2. Consequently, the model ascribes no role to financial innovation within M2 in determining the long-run level of demand for M1, merely a transient response to changes in R_{ma}. The importance of this finding is that we have modelled the relevant innovation, and still not found an effect, yet our equation experiences no problems with the three salient episodes noted above. Conversely, financial innovation within M1 is found to be important, especially in accounting for the M1 explosion period, as in the U.K. (see Hendry and Ericsson (1991)).

The inflation effects are large in the short run, but the mean lags are longer, at four to six quarters, which seems sensible for a transactions medium which doubles as an asset. Finally, the volatility effects are large in the short run, with both V_t and ΔSV_{t-1} influencing demand. We return shortly to an interpretation of their role.

Thus, all of the impact coefficients have interpretable signs with appropriate negative coefficients for responses to opportunity cost variables and feedbacks, and positive reactions to income, risk, and own interest rates. The parameterization corresponds to relatively orthogonal variables since only 11 of the 78 correlations between regressors exceed 0·5 in absolute value, and none exceeds 0·9. Conversely, eleven "t-values" exceed 5, and the smallest is 3·5. Moreover, of the 13 correlations between $\Delta(m-p)_t$ and its explanatory variables, all but one of the partial correlations exceeds the corresponding simple correlation, and two have the opposite sign, consistent with a structural interpretation of (18).

The value of $\hat{\sigma}$ of 0·38% entails a tight fit, as can be seen in Figure 5 (however, the unconditional standard deviation of Δm_t is just under 1%). Equally, an overall F-test on the validity of (18) against the GUM yields a value of less than unity. Finally, it is unsurprising that the design has achieved "insignificant" values for most of the diagnostic checks (e.g. the serial correlation statistics for up to fourth order), although it must be remembered that the t-value cutoff was high, thus "dumping" other lagged variables into

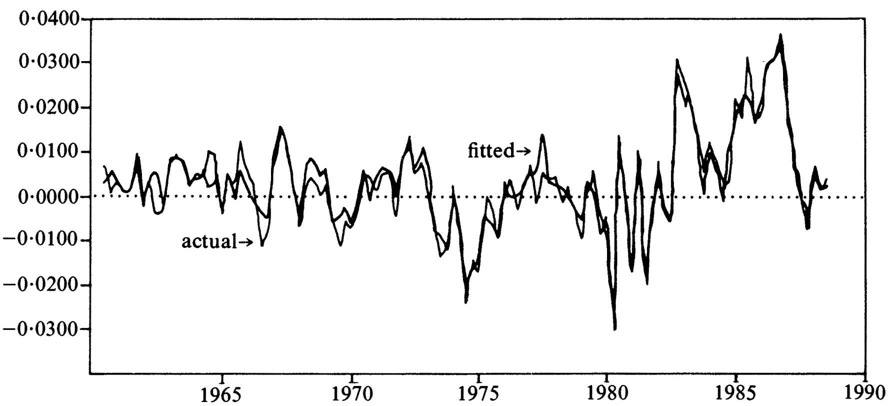

FIGURE 5

Fitted and actual values from (18) for $\Delta(m-p)$

the residual even if their coefficients were as large as twice their standard errors. In fact, there is some evidence of eighth-order autocorrelation, although that again might be an artifact of separate prior seasonal adjustment.

The most important features of (18) remaining to be tested are its constancy over time, and if that is not rejected, its ability to account for the predictive failures and successes of earlier models. These are the subject of Section VII. First we focus on a slightly more parsimonious representation intended to highlight the role of the risk measures. An auxiliary regression of AS_t on V_t and ΔSV_{t-1} is used to orthogonalize their effects:

$$\underset{(0{\cdot}04)}{AS_t = 0{\cdot}57\,V_t} + \underset{(4{\cdot}0)}{6{\cdot}4\Delta SV_{t-1}} \tag{21}$$

$$R^2 = 0{\cdot}68, \qquad \hat{\sigma} = 1{\cdot}17\%, \qquad F(2, 111) = 120{\cdot}2, \qquad DW = 0{\cdot}16.$$

Interpreting the residuals from (21) as a "risk-adjusted spread" AS_t^* yields:

$$AS_t^* = AS_t - 0{\cdot}57\,V_t - 6{\cdot}4\Delta SV_{t-1}.$$

On replacing AS_t by AS_t^* in (18), an F-test for eliminating the two variables (V_t, ΔSV_{t-1}) was not significant at the 5% level ($F_{99}^2 = 2{\cdot}7$) and the Schwarz criterion fell to −10.70. Imposing these restrictions produced the simplified model:

$$\begin{aligned}\Delta(m-p)_t = {} & \underset{(0{\cdot}020)}{0{\cdot}358} - \underset{(0{\cdot}098)}{0{\cdot}348\Delta_4(m-p)_{t-1}} - \underset{(0{\cdot}015)}{0{\cdot}254(m-p-\tfrac{1}{2}y)_{t-2}} - \underset{(0{\cdot}104)}{1{\cdot}428AS_t^*} + \underset{(0{\cdot}070)}{0{\cdot}370\Delta Ay_t} \\ & - \underset{(0{\cdot}063)}{0{\cdot}985AR_{1_t}} - \underset{(0{\cdot}049)}{0{\cdot}260\Delta R_{ma_t}} - \underset{(0{\cdot}129)}{1{\cdot}066\Delta_4 p_{t-1}} - \underset{(0{\cdot}046)}{0{\cdot}341\Delta\hat{p}_t} + \underset{(0{\cdot}051)}{0{\cdot}465R_{nsa_t}} \\ & - \underset{(0{\cdot}040)}{0{\cdot}148\Delta^2(m-p)_{t-4}} + \underset{(0{\cdot}003)}{0{\cdot}013D_t}\end{aligned} \tag{22}$$

$$R^2 = 0{\cdot}889, \qquad \hat{\sigma} = 0{\cdot}391\%, \qquad F(11, 101) = 73{\cdot}3, \qquad DW = 1{\cdot}79, \qquad SC = -10{\cdot}70.$$

This model is consistent with the role of V_t and ΔSV_{t-1} being purely to *risk adjust* AS_t, and we interpret AS_t^* as the "certainty equivalent" opportunity cost discussed above. Figure 6 shows the graph of AS_t^* together with AS_t for comparison: while similar, the

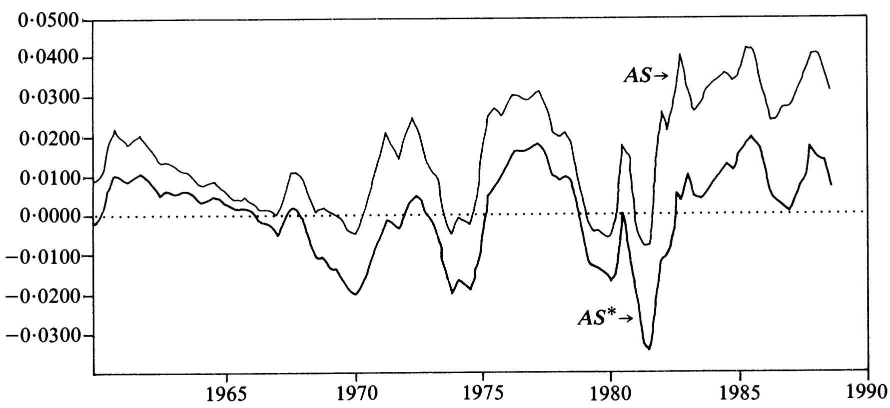

FIGURE 6

Time series of averages of spread and risk-adjusted spread: *AS*, *AS** times during which AS_t^* deviates most from AS_t are when earlier models experienced predictive difficulties.[12]

VI. ALTERNATIVE SPECIFICATIONS AND RELATIONS TO OTHER STUDIES

Table II summarizes a sample of other U.S. M1 demand studies: Judd and Scadding (1982) and Goldfeld and Sichel (1990) provide more comprehensive views. Most studies have an earlier starting point than the present analysis, so we cannot claim strict comparability. Nevertheless, with two exceptions (lines 4D, 5), all have equation standard errors

TABLE II

U.S. real M1 demand studies

Investigator/source	Dates	Standard error	Constancy	Comments
1. Goldfeld (1973)	52(2)-72(2)	0·43%	rejects 74(1)-76(2)	Goldfeld (1976)
2. Garcia-Pak (1979)	52(2)-76(2)	0·63%		
3. Rose (1985)	52(2)-71(4)	0·48%	rejects 81(4)	Inflation significant
4. Gordon (1984)				
model D	56(3)-72(4)	0·38%	rejects 74-76	Inflation
	56(3)-83(1)	0·50%	rejects 80-83	
model F	56(3)-72(4)	0·43%	accepts	Error correction model No coefficient significant (1958-1972)
	56(3)-83(1)	0·48%		Inflation significant
5. McAleer-Pagan-Volker (1985)	52(2)-73(4)	0·31%	accept over sample reject out-of sample	Inflation significant
6. Simpson-Porter (1980)	55(1)-74(2)	0·46%-0·47%	reject 1981(4) (Offenbacher-Porter (1982)	Cash management ratchet significant
	55(1)-80(2)	0·52%-0·59%		
7. Cooley-Leroy (1981)	52(2)-78(4)	2·80%	unreported	Static structure; large error
8. Baseline (18)	60(3)-88(3)	0·38%	accept	Error-correction, inflation, and bond volatility significant
	60(3)-72(4)	0·38%		

12. On the old BHS data, we obtained the following estimates for the revised specification over the sample 1960(3) to 1984(2) with 6 outside sample forecasts:

$$\Delta(m-p)_t = \underset{(0{\cdot}055)}{-0{\cdot}712} - \underset{(0{\cdot}134)}{0{\cdot}376}\Delta_4(m-p)_{t-1} - \underset{(0{\cdot}022)}{0{\cdot}290}(m-p-\tfrac{1}{2}y)_{t-2} - \underset{(0{\cdot}140)}{1{\cdot}476}AS_t^* - \underset{(0{\cdot}089)}{1{\cdot}045}AR_{1_t}$$

$$-\underset{(0{\cdot}049)}{0{\cdot}215}\Delta Rma_t - \underset{(0{\cdot}165)}{0{\cdot}860}\Delta_4 p_{t-1} - \underset{(0{\cdot}054)}{0{\cdot}324}\Delta\hat{p}_t + \underset{(0{\cdot}077)}{0{\cdot}436}\Delta Ay_t + \underset{(0{\cdot}110)}{0{\cdot}429}Rnsa_t$$

$$-\underset{(0{\cdot}044)}{0{\cdot}119}\Delta^2(m-p)_{t-4} + \underset{(0{\cdot}003)}{0{\cdot}013}D_t$$

$R^2=0{\cdot}856$, $\hat{\sigma}=0{\cdot}380\%$, $F(11,80)=45{\cdot}49$, $DW=1{\cdot}63$, $SC=-10{\cdot}71$, Mean $=0{\cdot}17\%$

S.D. $=0{\cdot}96\%$, $\eta_1(6,78)=1{\cdot}47$, $\eta_2(6,84)=2{\cdot}57$, $\eta_3(22,61){\cdot}54$

$\xi_5(2)=0{\cdot}60$, $\eta_6(4,76)=0{\cdot}90$, $\eta_7(1,83)=0{\cdot}06$.

This is close to (22) and fits almost as well as the model reported in (1986) which had $\hat{\sigma}=0{\cdot}378\%$. Thus, the data revisions and our re-specification of volatility have mainly altered the dynamic specification.

larger by a tenth to a half. Of the exceptions, Gordon (1984) in 1956–1972, specification D repeatedly rejects parameter constancy, and McAleer, Pagan and Volker (1985) is not estimated beyond 1973. Those estimated through the 1970's to early 1980's reject constancy somewhere in that period with the exception of Gordon's (1984) model F, for which no coefficient estimate is significant 1956–1972. The Simpson and Porter (1980) study attempts to quantify the impetus to financial innovation by introducing a uni-directional ("ratchet") variable driven by new peaks in interest rates. The model thus derived tracks the missing money period, but founders in the early 1980's as the ratchet variable continues to increase (with new peaks in interest rates), hence predicting a corresponding downward adjustment in real money demand which was not realized (see Offenbacher and Porter (1982)). Rose (1985) uses an error-correction model and finds inflation significant, but eventually rejects constancy. Cooley and Leroy (1981) use a static model resulting in an equation standard error several times as large as in other models.

The most common alternatives to the error-correction treatment of dynamics are partial adjustment models in which the dynamic structure is summarized by a lagged dependent variable, perhaps allowing for residual autocorrelation. The analytical properties of such equations are discussed in Hendry *et al.* (1984), but to characterize this alternative empirically, we investigate a partial adjustment model. It results in dramatically less precise estimates: standard errors go up by about half to 0·576%. There is also a degeneration in parameter constancy tests, suggesting that estimates in some (more volatile) time periods are more dependent than others on a flexible dynamic specification.

The partial adjustment model corresponding to (18) is presented in Table III; its comparison with the error-correction model points up the strengths of the latter. On goodness-of-fit, the standard error of the partial adjustment model is almost 50% greater than (18). Split-sample covariance tests of parameter constancy over 1960(3)–1974(2) versus 1974(3)–1988(3) and of 1960(3)–1979(3) versus 1979(4)–1984(2) yield $F^9_{93} = 6.01$ and $F^9_{76} = 4{\cdot}42$, so both reject constancy at the 1% level, indicating a structural breakdown

TABLE III

Alternative models

Partial adjustment model			Goldfeld (1976) model		
	1960(3)–1988(3)			1960(3)–1988(3)	
	Coeff.	St. Error		Coeff.	St. Error
CONSTANT	0·277	0·032	CONSTANT	0·035	0·077
$(m-p)_{t-1}$	0·808	0·021	y_t	0·041	0·019
Δp_t	−0·901	0·141	$\ln Rp_t$	0·005	0·023
y_t	0·090	0·009	$\ln R_{1_t}$	−0·012	0·005
S_t	−1·158	0·220	$(m_{t-1}-p_t)$	0·950	0·024
V_t	0·549	0·104	D_t	0·016	0·004
SV_t	7·361	5·002	α_{-1}	0·485	0·094
R_{1_t}	−0·713	0·124			
Rma_t	−0·057	0·073			
$Rnsa_t$	0·354	0·087			
D_t	0·016	0·004			
T	113		T	112	
K	11		K	7	
$\hat{\sigma}$	0·576%		$\hat{\sigma}$	0·662%	
R^2	0·996		$\xi_5(2)$	5·19	
η_1	4·74				
$\xi_5(2)$	0·31				

when the dynamic structure is inadequate. The mean lag response to income (or any other variable) is estimated at around 4-5 quarters, but there is fourth-order residual serial correlation: even correcting the latter leaves $\hat{\sigma} = 0{\cdot}55\%$. Nevertheless, the partial adjustment model empirically agrees with (18) on the significance of the distinctive theoretical approach undertaken. In particular, the yield and risk to long-term bond holding (S and V), inflation, and learning-adjusted M1-own yield are found to be significant determinants of M1 demand in Table III.

Since the present study is estimated over a different sample from Goldfeld (1976), and on revised data, it is appropriate to re-estimate the demand function of that study on the same sample for purposes of comparison. The specification is:

$$(m_t - p_t) = \beta_0 + \beta_1 \ln \mathrm{GNP}_t - \beta_2 \ln R_{pt} - \beta_3 \ln R_{1_t} + \beta_4(m_{t-1} - p_t) + \beta_5 D_t. \tag{23}$$

The model is estimated with a first-order autoregressive error, thus imposing a common factor constraint (see Hendry and Mizon (1978)): that constraint is strongly rejected on the data sample here. Interestingly, so is parameter constancy within the sample Goldfeld (1973) used, confirming Koskela and Viren (1984) who also reject his specification on both grounds over earlier sample periods. Results are shown for comparison purposes in Table III. On the present vintage of data revisions, the Chow test does not reject the Goldfeld model over *missing money*, but the predictive failure statistic $\xi_9(H)/H$ indicates considerable inaccuracy (see Table VI).

The next group of alternative specifications involves the deletion of variables important to the study, namely V, SV and Δp. Again, there is a significant deterioration in fit and, for deletion of V and Δp, in parameter constancy. These comparisons are shown in Section VII and confirm the appropriateness of the choice of variables.

VII. SUB-PERIOD PARAMETER CONSTANCY

Most models of U.S. M1 demand fail tests of parameter constancy over one or more of the following sub-periods: *missing money*, 1974(1)–1976(2); *great velocity decline*, 1982(1)–1983(2); *M1 explosion*, 1985(1)–1986(4). It is our view that the rejection of parameter constancy reflects model mis-specification rather than shifts in the demand function since the baseline model, (18) in Section V, accepts constancy in each of these sub-periods. We conclude that accounting for dynamic structure, long-term bond risk and return, financial innovation, and inflation are sufficient for a constant parameter M1 demand function over these difficult sub-periods.

The sub-sample periods displayed in Tables IV and V correspond to distinctive periods in recent monetary history. The end of the first sample in 1972(4) corresponds approximately to the end of the period over which Goldfeld's (1973) classic study was based. The period ending in 1976(2) concludes the period of structural breakdown of most money demand models, namely the period of missing money (Goldfeld (1976): see Section VIIa). The sample ending in 1979(3) completes the period prior to the introduction of the Federal Reserve System's New Operating Procedures. The next sample to 1985(4) includes the period of those procedures, one quarter of credit controls, and an unprecedented decline in monetary velocity: this also coincides with the longest test sample in BHS (see VIIb). The last sample covers the M1 explosion, continuing up to 1988(3) (see VIIc). The entire sample is characterized by financial institutional innovation, changing the composition, liquidity and yield of financial instruments competitive with, or internal to, M1, and the oft-cited breakdown of existing equations.

TABLE IV

Baseline model with $\Delta(m-p)$ as dependent variable

	1960(3) 1985(4)	1960(3) 1972(4)	1973(1) 1985(4)	1960(3) 1976(2)	1960(3) 1979(3)	1960(3) 1988(3)
CONSTANT	0·336	0·409	0·345	0·380	0·316	0·352
	(0·030)	(0·082)	(0·037)	(0·057)	(0·044)	(0·020)
$\Delta_4(m-p)_{t-1}$	−0·351	−0·185	−0·539	−0·212	−0·256	−0·334
	(0·135)	(0·240)	(0·252)	(0·181)	(0·172)	(0·097)
$\Delta^2(m-p)_{t-4}$	−0·149	−0·227	−0·126	−0·192	−0·134	−0·156
	(0·043)	(0·100)	(0·053)	(0·074)	(0·067)	(0·039)
$\Delta\hat{p}_t$	−0·341	−0·312	−0·443	−0·343	−0·328	−0·330
	(0·051)	(0·084)	(0·083)	(0·061)	(0·058)	(0·046)
$\Delta_4 p_{t-1}$	−1·053	−1·667	−1·337	−1·124	−0·849	−1·097
	(0·172)	(0·555)	(0·440)	(0·314)	(0·268)	(0·132)
ΔAy_t	0·407	0·469	0·338	0·439	0·442	0·395
	(0·077)	(0·132)	(0·104)	(0·101)	(0·090)	(0·070)
$(m-p-\frac{1}{2}y)_{t-2}$	−0·238	−0·296	−0·239	−0·274	−0·223	−0·249
	(0·021)	(0·058)	(0·026)	(0·041)	(0·031)	(0·015)
AS_t	−1·347	−1·437	−1·388	−1·308	−1·333	−1·409
	(0·143)	(0·526)	(0·188)	(0·261)	(0·237)	(0·104)
V_t	0·823	0·897	0·886	0·754	0·772	0·859
	(0·095)	(0·310)	(0·150)	(0·126)	(0·123)	(0·079)
ΔSV_{t-1}	11·680	14·126	10·851	13·430	12·792	11·683
	(1·613)	(7·449)	(1·847)	(4·593)	(4·447)	(1·490)
AR_{1t}	−0·931	−0·908	−0·958	−0·922	−0·918	−0·973
	(0·087)	(0·310)	(0·118)	(0·195)	(0·182)	(0·063)
ΔRma_t	−0·246	−0·220	−0·190	0·128	−0·201	−0·255
	(0·053)	(0·649)	(0·061)	(0·306)	(0·189)	(0·049)
$Rsna_t$	0·445		0·430			0·435
	(0·072)		(0·088)			(0·055)
D_t	0·013		0·013			0·013
	(0·003)		(0·003)			(0·003)
T	102	50	52	64	77	113
K	14	12	14	12	12	14
$\hat{\sigma}$	0·400%	0·390%	0·418%	0·371%	0·381%	0·385%
R^2	0·864	0·703	0·920	0·824	0·780	0·894
η_1	$\eta_1(4,84)=0{\cdot}22$	$\eta_1(4,34)=0{\cdot}64$	$\eta_1(4,34)=0{\cdot}60$	$\eta_1(4,48)=0{\cdot}61$	$\eta_1(4,95)=0{\cdot}88$	$\eta_1(4,95)=0{\cdot}12$
$\xi_5(2)$	0·70	0·14	0·93	0·17	0·88	0·45

Parameter constancy tests of the baseline model (18) and the risk-adjusted yield spread model (22) are as follows:

		Baseline model (18)	Risk-adjusted spread model (22)
Missing money Chow test	1960(3)-1973(4) vs. 1960(3)-1976(2)	$F^{10}_{42}=0{\cdot}67$	$F^{10}_{44}=0{\cdot}76$
New operating procedures Chow test	1960(3)-1979(3) vs. 1960(3)-1982(3)	$F^{10}_{65}=0{\cdot}71$	$F^{10}_{67}=1{\cdot}38$
Great velocity decline Chow test	1960(3)-1981(4) vs. 1960(3)-1983(2)	$F^{6}_{72}=0{\cdot}87$	$F^{6}_{74}=2{\cdot}27$
M1 explosion Chow test	1960(3)-1984(4) vs. 1960(3)-1986(4)	$F^{8}_{84}=1{\cdot}53$	$F^{8}_{86}=1{\cdot}13$
Split-sample covariance test	1960(3)-1974(2) vs. 1974(3)-1988(3)	$F^{12}_{87}=1{\cdot}18$	$F^{10}_{91}=1{\cdot}29$

TABLE V

Baseline model with $\Delta(m-p)$ as the dependent variable and in which risk is incorporated in the spread

	1960(3) 1985(4)	1960(3) 1972(4)	1973(1) 1985(4)	1960(3) 1976(2)	1960(3) 1979(3)	1960(3) 1988(3)
CONSTANT	0·359	0·416	0·361	0·388	0·323	0·358
	(0·028)	(0·080)	(0·035)	(0·051)	(0·040)	(0·021)
$\Delta_4(m-p)_{t-1}$	−0·416	−0·288	−0·672	−0·261	−0·304	−0·348
	(0·133)	(0·210)	(0·240)	(0·172)	(0·163)	(0·098)
$\Delta^2(m-p)_{t-4}$	−0·134	−0·193	−0·108	−0·193	−0·134	−0·148
	(0·043)	(0·091)	(0·052)	(0·071)	(0·065)	(0·040)
$\Delta\hat{p}_t$	−0·344	−0·287	−0·469	−0·332	−0·320	−0·341
	(0·052)	(0·077)	(0·081)	(0·057)	(0·053)	(0·046)
$\Delta_4 p_{t-1}$	−1·099	−1·460	−1·554	−1·093	−0·822	−1·066
	(0·170)	(0·449)	(0·421)	(0·277)	(0·212)	(0·129)
ΔAy_t	0·393	0·476	0·327	0·441	0·450	0·370
	(0·078)	(0·126)	(0·104)	(0·100)	(0·089)	(0·070)
$(m-p-\frac{1}{2}y)_{t-2}$	−0·254	−0·300	−0·248	−0·279	−0·228	−0·253
	(0·020)	(0·057)	(0·025)	(0·038)	(0·029)	(0·015)
AS^*_t	−1·437	−1·470	−1·438	−1·363	−1·384	−1·428
	(0·140)	(0·514)	(0·180)	(0·212)	(0·202)	(0·105)
AR_{1t}	−0·985	−0·975	−0·965	−0·976	−0·965	−0·985
	(0·084)	(0·288)	(0·106)	(0·147)	(0·143)	(0·063)
ΔRma_t	−0·260	−0·230	−0·194	0·081	−0·224	−0·260
	(0·053)	(0·621)	(0·061)	(0·296)	(0·185)	(0·049)
$Rsna_t$	0·491		0·452			0·465
	(0·069)		(0·088)			(0·051)
D_t	0·013		0·013			0·013
	(0·003)		(0·003)			(0·003)
T	102	50	52	64	77	113
K	12	10	12	10	10	12
$\hat{\sigma}$	0·407%	0·385%	0·421%	0·368%	0·377%	0·391%
R^2	0·857	0·696	0·915	0·820	0·777	0·889
η_1	0·35	0·79	0·38	0·84	1·07	0·35
$\xi_5(2)$	0·52	0·28	0·99	0·20	0·80	0·37

None of the tests on (18) rejects constancy at the 5% level.[13] The sub-period variation of $\hat{\sigma}$ between 0·37% and 0·42% is small and any large unaccounted-for source of deviation in a single observation can be enough to cause rejection in the parameter constancy test (e.g. if the credit control dummy is deleted, constancy is rejected at the 1% level over 1980(1)–1981(4)). Since the relationship is known to have "broken down" in other formulations, the constancy tests are demanding. Figure 7 shows the 1-step ahead conditional predictions from (18), and the associated 95% confidence intervals, together with the realized values over the 20 periods 1983(4)–1988(3) in terms of $(m-p)_t$. The

13. The tests of parameter constancy used are the covariance test (often called "Chow"):

$$\eta_2^* = F^{K_1+K_2-K_3}_{T-K_1-K_2} = \frac{RSS_T-(RSS_{T_1}+RSS_{T_2})}{RSS_{T_1}+RSS_{T_2}} \cdot \frac{T-K_1-K_2}{K_1+K_2-K_3} \quad \text{for } T_1>K_1,\ T_2>K_2;$$

and the Chow test:

$$\eta_2 = F^{T-T_1-(K-K_1)}_{T_1-K_1} = \left(\frac{RSS}{RSS_1}-1\right) \cdot \frac{T_1-K_1}{T-T_1-(K-K_1)}.$$

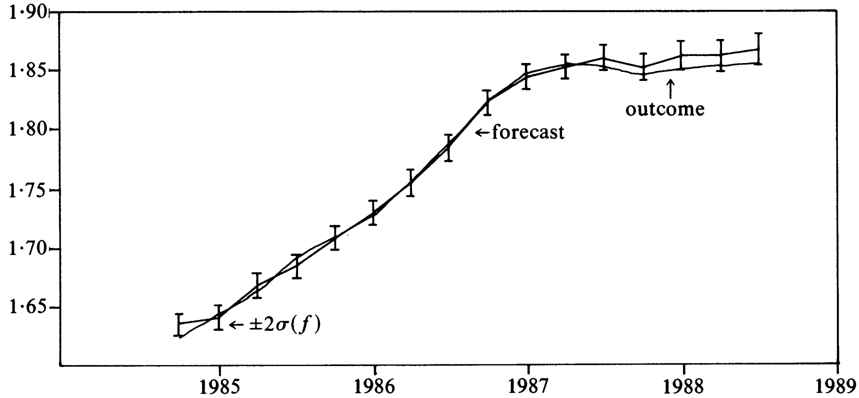

FIGURE 7
One-step ahead forecasts and outcomes for $(m-p)$ based on (18), with ± 2 forecast standard errors around the forecast $[\sigma(f)]$

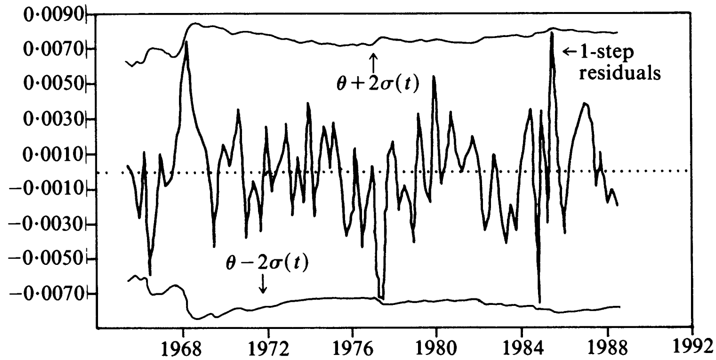

FIGURE 8
One-step residuals from (22) with ± 2 equation standard errors (σ) over 1965(4)–1988(3)

constancy can be seen graphically from the recursive residuals in Figure 8, shown together with $0 \pm 2\hat{\sigma}$ for every increasing sub-sample.[14]

Turn now to the model with the risk-adjusted yield spread (22). The theoretical model of Section II suggests that the risk and return of long term bonds should enter the demand for M1. They do so in both (18) and (22), in (22) entering in the simplified form of a risk discount on the yield spread of bonds over bills. The simplified form in (22) emphasizes the idea that the only way bond price volatility enters money holding is through wealth holder portfolio behaviour. In contrast, in the more complex model (18), a skeptical view of the bond risk variables might be to construe them as acting as dummies for periods of known difficulty in other models. Hence one test of the soundness of the approach here is consistency of (18) and (22). The coefficients on the yield spread variables (AS_t in (18), AS_t^* in (22)) are virtually identical. The only significant discrepancy

14. Equation (22) was used to compute recursive estimates, with the coefficient of ΔAy_t restricted to $\frac{1}{2}$ and with the coefficient of the dummy imposed at its full sample estimate: an F-test of the first restriction yields $F(1, 101) = 3{\cdot}2$ (not significant at the 5% level). Two early values of R_{nsa} were set to $\pm 0{\cdot}00001$ to initialize the recursions. See Brown, Durbin and Evans (1975), Dufour (1982) and Hendry (1989) on these recursive procedures.

occurs in the great velocity decline where the Chow test on (22) is (barely) significant at the 5% level. Even for a constant parameter model, one out of a dozen F tests might be significant at the 5% level. Indeed, the constancy of (22) holds for all break-point Chow (1960) tests; and over the problem periods, the coefficients of (18) and (22) are closely similar as Tables IV and V show.

For this money demand function, the missing money was never gone. Constancy similarly holds over the great velocity decline (with mild difficulty for (22)), which has troubled other equations, and over the M1 explosion, where real M1 increased by more than $50\hat{\sigma}$. Nor is tracking M1 after the introduction of the New Operating Procedures a major difficulty, so the model does not seem susceptible to policy regime shifts either. Thus, we can now consider the possible causes of the failures of earlier models.

A. *The case of the missing money: an arrest is made*

The period 1974(1)–1976(2) is characterized by structural breakdown in models of M1 demand. Typically, models over-predicted demand by 7% to 12% with the most widely known model over-predicting by 8·7% (Goldfeld (1976)). Attempts to analyse and explain the apparent fall in M1 demand focused on institutional change, financial innovation, and overnight conversion of demand deposits to interest-bearing form (Judd and Scadding (1982), Garcia and Pak (1979)). Goldfeld and Sichel (1990) conclude that the problem remains unresolved. The constancy over that sample of the present model provides a solution. We have already shown that M2 financial innovation is not the explanation.

Omission of the risk and return to long-term assets, measured as bond holding-period volatility and long-term bond yields, is a mis-specification that accounts for the breakdown of the Goldfeld standard specification in the missing-money period. We demonstrate this contention as follows: both baseline models (18) and (22) are constant; but if the model is estimated excluding bond volatility variables, then constancy is rejected. These results are reported in Table VI. This is a smoking gun: one of the usual suspects, interest rates, in particular long-term rates with an allowance for risk, took the missing money. Figures 9 and 10 show one-step conditional predictions and 95% confidence intervals from (22) and the Goldfeld specification respectively, over the period 1974(1)–1976(2).

B. *The great velocity decline*

Over 1982 and the first half of 1983, many models of M1 demand recorded systematic under-prediction errors: the great velocity decline. The ratio of GNP at an annual rate to the money stock fell at an historically unprecedented rate. The decline of velocity in the 1980s was more dramatic than the apparent instability of the mid-1970s. Since this fall was unpredicted by most models, the prevailing (though not unanimous) view was that there had been a shift in the money demand function. Models reported in Federal Reserve Bank of San Francisco (1983) have six-quarter (1982(1)–1983(2)) fitted errors of −2·8% to −8·3%. Other models in use, however, over-predict (see Judd and Motley (1984)).

Omission of bond volatility and the long-term bond yield from money demand functions is a mis-specification that accounts for structural breakdown in 1982(1)–1983(2). The baseline model (18) is constant over that period, yet constancy is rejected if inflation and volatility variables are omitted from it.[15] These results are reported in Table VII.

15. The yields on NOW and SuperNOW accounts also enter the model but are not essential in this interval.

TABLE VI

Missing money models

	1960(3)–1973(4)		1960(3)–1976(2)		Forecast tests
Baseline (18)	T	54	T	64	$\eta_2(10, 42) = 0{\cdot}67$
	K	12	K	12	$\chi^2_{10}/10 = 1{\cdot}89$
	$\hat{\sigma}$	0·384%	$\hat{\sigma}$	0·371%	
Baseline with volatility deleted	T	54	T	64	$\eta_2(10, 44) = 2{\cdot}78$
	K	10	K	10	$\chi^2_{10}/10 = 15{\cdot}22$
	$\hat{\sigma}$	0·424%	$\hat{\sigma}$	0·489%	
Baseline with inflation deleted	T	54	T	64	$\eta_2(10, 44) = 1{\cdot}51$
	K	10	K	10	$\chi^2_{10}/10 = 5{\cdot}75$
	$\hat{\sigma}$	0·461%	$\hat{\sigma}$	0·483%	
Baseline without any learning adjustment	T	54	T	64	$\eta_2(10, 42) = 0{\cdot}68$
	K	12	K	12	$\chi^2_{10}/10 = 2{\cdot}67$
	$\hat{\sigma}$	0·383%	$\hat{\sigma}$	0·371%	
Baseline which uses *Rp* rather than *Rma*	T	54	T	64	$\eta_2(10, 42) = 0{\cdot}76$
	K	12	K	12	$\chi^2_{10}/10 = 1{\cdot}96$
	$\hat{\sigma}$	0·380%	$\hat{\sigma}$	0·371%	
Partial Adjustment Model	T	54	T	64	$\eta_2(10, 45) = 0{\cdot}68$
	K	9	K	9	$\chi^2_{10}/10 = 1{\cdot}91$
	$\hat{\sigma}$	0·463%	$\hat{\sigma}$	0·449%	
Goldfeld Model	T	53	T	63	$\eta_2(10, 47) = 1{\cdot}68$
	K	6	K	6	$\chi^2_{10}/10 = 7{\cdot}57$
	$\hat{\sigma}$	0·434%	$\hat{\sigma}$	0·449%	
Risk-adjusted spread model (22)	T	54	T	64	$\eta_2(10, 44) = 0{\cdot}76$
	K	10	K	10	$\chi^2_{10}/10 = 1{\cdot}33$
	$\hat{\sigma}$	0·377%	$\hat{\sigma}$	0·368%	

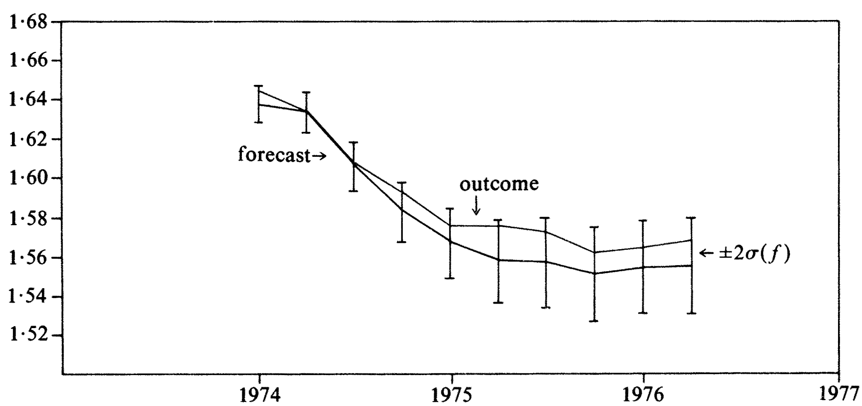

FIGURE 9

One-step ahead forecasts and outcomes for $(m-p)$ based on (18) with ±2 forecast standard errors $[\sigma(f)]$ around the forecast

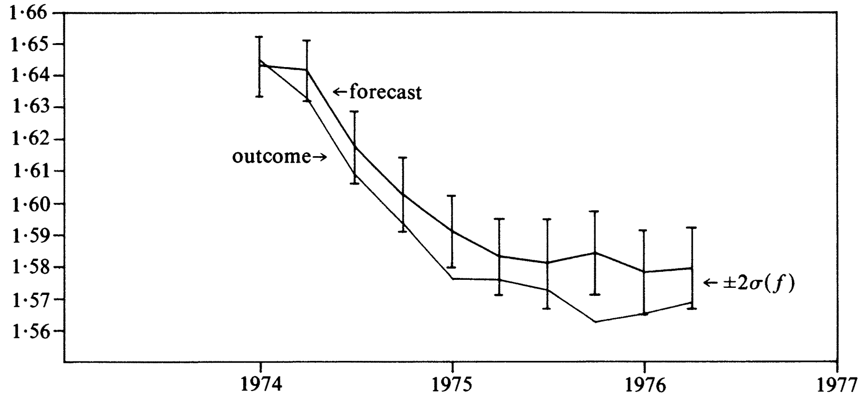

FIGURE 10

One-step ahead forecasts and outcomes for $(m-p)$ based on the Goldfeld model with ±2 forecast standard errors.

TABLE VII

Great velocity decline models

	1960(3)-1981(4)		1960(3)-1983(2)		Forecast tests
Baseline (18)	T	86	T	92	$\eta_2(6, 72) = 0{\cdot}87$
	K	14	K	14	$\chi^2_6/6 = 13{\cdot}96$
	$\hat{\sigma}$	0·376%	$\hat{\sigma}$	0·374%	
Baseline with volatility deleted	T	86	T	92	$\eta_2(6, 74) = 5{\cdot}09$
	K	12	K	12	$\chi^2_6/6 = 20{\cdot}27$
	$\hat{\sigma}$	0·486%	$\hat{\sigma}$	0·556%	
Baseline with inflation deleted	T	86	T	92	$\eta_2(6, 74) = 3{\cdot}61$
	K	12	K	12	$\chi^2_6/6 = 8825{\cdot}88$
	$\hat{\sigma}$	0·514%	$\hat{\sigma}$	0·562%	
Baseline with interest on checkables deleted	T	86	T	92	$\eta_2(6, 73) = 2{\cdot}61$
	K	13	K	13	$\chi^2_6/6 = 6{\cdot}33$
	$\hat{\sigma}$	0·374%	$\hat{\sigma}$	0·396%	
Baseline without any learning adjustment	T	86	T	92	$\eta_2(6, 72) = 1{\cdot}32$
	K	14	K	14	$\chi^2_6/6 = 15{\cdot}46$
	$\hat{\sigma}$	0·381%	$\hat{\sigma}$	0·385%	
Baseline which uses *Rp* rather than *Rma*	T	86	T	92	$\eta_2(6, 72) = 1{\cdot}76$
	K	14	K	14	$\chi^2_6/6 = 1608{\cdot}43$
	$\hat{\sigma}$	0·402%	$\hat{\sigma}$	0·414%	
Partial Adjustment	T	86	T	92	$\eta_2(6, 75) = 2{\cdot}76$
	K	11	K	11	$\chi^2_6/6 = 997{\cdot}10$
	$\hat{\sigma}$	0·494%	$\hat{\sigma}$	0·525%	
Goldfeld Model	T	85	T	91	$\eta_2(6, 78) = 6{\cdot}46$
	K	7	K	7	$\chi^2_6/6 = 7{\cdot}18$
	$\hat{\sigma}$	0·497%	$\hat{\sigma}$	0·577%	
Risk-adjusted spread model (22)	T	86	T	92	$\eta_2(6, 74) = 2{\cdot}27$
	K	12	K	12	$\chi^2_6/6 = 371{\cdot}83$
	$\hat{\sigma}$	0·375%	$\hat{\sigma}$	0·392%	

The period 1979-1983, spanning the reversal in the trend growth in velocity, was of signal importance and difficulty for monetary policy, which, in this apparently unstable period, was predicated on predictability of money demand. In October 1979, the Federal Reserve System, as an anti-inflation measure, adopted the New Operating Procedures which emphasized contemporaneous control of the monetary base in order to foster steady and reduced growth of the base so as to target corresponding growth in the monetary aggregates, particularly M1. The policy explicitly disavowed any role for interest rate stabilization, and interest rate volatility (measured by V_q) tripled over the course of a year.

Goldfeld and Sichel (1990) comment:

> "It is perhaps ironic that ... [money demand instability] roughly coincided with the adoption by a number of central banks of policies aimed at targeting monetary aggregates. Some have argued that this association is more than mere coincidence."

The full application of the New Operating Procedures was interrupted in 1980(2) by one quarter of credit control, generating a short recession. 1981-1982 then had the severest recession in the United States since the 1930's. In August 1982, noting that the relationship of money to output had become unreliable, the Fed announced that it would no longer target M1, and a regime of interest rate targeting followed.

The confusion over monetary policy in 1982 (and the decline in velocity) represented the unforeseen consequence of the policy change undertaken in 1979. However, the effects were in fact predictable as the results in Table VII show. The demand function estimated through 1979(3) (prior to the start of the New Operating Procedures) tracks consistently through 1982(3). The value of the coefficient estimates of variables representing the risk to holding long-term bonds do not change significantly and remain statistically significant throughout the period, even for samples finishing as early as 1972. (18) is stable over the great velocity decline and (22) is essentially so: Chow tests for (22) of the five and seven quarter periods 1982(1)-1983(1) and 1982(1)-1983(3) are not significant at the 5% level.

We conclude that the apparent shifts in money demand in 1982 represented the foreseeable consequence on money demand of interest rate volatility: that volatility was itself the result of central bank policy and the confusion in monetary policy was fully avoidable.

C. The M1 explosion of 1985-1986

Goldfeld and Sichel (1990) note that money demand models typically do not display parameter constancy after 1979(3). Even when artificially adjusted to display constancy through 1984(4) (using dummies), constancy fails again in 1985 and 1986:

> "this is perhaps hardly surprising when one notes that in 1985 and 1986, M1 grew at 12·1% and 15·3%, respectively, rates far exceeding typical ones, either absolutely or relative to the rate of change in real GNP."

The essential reasons for the increased demand are clear in the present model: declining bond and bill interest rates, steady low inflation, and adjustment by wealth holders to the availability of NOW and SuperNOW accounts.

Results of the baseline model and alternatives are presented in Table VIII. The baseline model experiences a mild deterioration in fit, but easily accepts constancy over the period. Alternative specifications omitting NOW and SuperNOW accounts, or omitting the gradual learning process for wealth holders to adapt to them, strongly reject

TABLE VIII

M1 explosion models

	1960(3)-1984(4)		1960(3)-1986(4)		Forecast tests
Baseline (18)	T	98	T	106	$\eta_2(8, 84) = 1{\cdot}53$
	K	14	K	14	$\chi_8^2/8 = 3{\cdot}78$
	$\hat{\sigma}$	0·386%	$\hat{\sigma}$	0·394%	
Baseline with volatility deleted	T	98	T	106	$\eta_2(8, 86) = 2{\cdot}20$
	K	12	K	12	$\chi_8^2/8 = 8{\cdot}52$
	$\hat{\sigma}$	0·578%	$\hat{\sigma}$	0·607%	
Baseline with inflation deleted	T	98	T	106	$\eta_2(8, 86) = 0{\cdot}62$
	K	12	K	12	$\chi_8^2/8 = 0{\cdot}69$
	$\hat{\sigma}$	0·608%	$\hat{\sigma}$	0·598%	
Baseline with interest on checkables deleted	T	98	T	106	$\eta_2(8, 85) = 7{\cdot}00$
	K	13	K	13	$\chi_8^2/8 = 13{\cdot}51$
	$\hat{\sigma}$	0·400%	$\hat{\sigma}$	0·492%	
Baseline without any learning adjustment	T	98	T	106	$\eta_2(8, 84) = 6{\cdot}75$
	K	14	K	14	$\chi_8^2/8 = 13{\cdot}56$
	$\hat{\sigma}$	0·388%	$\hat{\sigma}$	0·475%	
Baseline which uses *Rp* rather than *Rma*	T	98	T	106	$\eta_2(8, 84) = 1{\cdot}90$
	K	14	K	14	$\chi_8^2/8 = 5{\cdot}60$
	$\hat{\sigma}$	0·423%	$\hat{\sigma}$	0·439%	
Partial Adjustment Model	T	98	T	106	$\eta_2(8, 87) = 2{\cdot}21$
	K	11	K	11	$\chi_8^2/8 = 4{\cdot}55$
	$\hat{\sigma}$	0·534%	$\hat{\sigma}$	0·560%	
Goldfeld Model	T	97	T	105	$\eta_2(8, 90) = 1{\cdot}82$
	K	7	K	7	$\chi_8^2/8 = 2{\cdot}63$
	$\hat{\sigma}$	0·588%	$\hat{\sigma}$	0·607%	
Risk-adjusted spread model (22)	T	98	T	106	$\eta_2(8, 86) = 1{\cdot}13$
	K	12	K	12	$\chi_8^2/8 = 1{\cdot}71$
	$\hat{\sigma}$	0·400%	$\hat{\sigma}$	0·402%	

constancy. Alternative specifications omitting dynamic structure, inflation, or risk and return to long-term assets accept parameter constancy but with an increase in standard error of about 50%. Hence, the present model deals successfully with the difficulty posed by Goldfeld and Sichel (1990) during the M1 explosion.

D. *Post sample*

Inasmuch as the baseline model was constructed and estimated over 1960(3)–1985(4) in BHS, it represents a rationalization, rather than a forecast of that period. One test of the model then is post-sample performance on the period 1986(1)–1988(3).

Results are reported in Table IX. Surprisingly, it turns out that this was *not* a difficult period to predict and parameter constancy is easily accepted for (18) or (22). Alternative specifications also accept constancy but with perceptibly greater difficulty and inferior fit. The Goldfeld (1976) model is rejected, however. We conclude that the baseline model (18) remains constant post-sample.

E. *An $I(0)$ reformulation*

The final transformation of the model is to reduce all variables to $I(0)$ so that standard inference procedures apply to tests (see Hendry and Mizon, 1990). Define the error

TABLE IX

Post sample forecasting

	1960(3)-1985(4)		1960(3)-1988(3)		Forecast tests
Baseline (18)	T K $\hat{\sigma}$	102 14 0·400%	T K $\hat{\sigma}$	113 14 0·385%	$\eta_2(11, 88) = 0{\cdot}33$ $\chi^2_{11}/11 = 0{\cdot}44$
Baseline with volatility deleted	T K $\hat{\sigma}$	102 12 0·591%	T K $\hat{\sigma}$	113 12 0·611%	$\eta_2(11, 90) = 1{\cdot}63$ $\chi^2_{11}/11 = 2{\cdot}55$
Baseline with inflation deleted	T K $\hat{\sigma}$	102 12 0·610%	T K $\hat{\sigma}$	113 12 0·621%	$\eta_2(11, 90) = 1{\cdot}35$ $\chi^2_{11}/11 = 2{\cdot}86$
Baseline with interest on checkables deleted	T K $\hat{\sigma}$	102 13 0·476%	T K $\hat{\sigma}$	113 13 0·490%	$\eta_2(11, 89) = 1{\cdot}52$ $\chi^2_{11}/11 = 2{\cdot}60$
Baseline without any learning adjustment	T K $\hat{\sigma}$	102 14 0·461%	T K $\hat{\sigma}$	113 14 0·467%	$\eta_2(11, 88) = 1{\cdot}26$ $\chi^2_{11}/11 = 2{\cdot}59$
Baseline which uses *Rp* rather than *Rma*	T K $\hat{\sigma}$	102 14 0·444%	T K $\hat{\sigma}$	113 14 0·432%	$\eta_2(11, 88) = 0{\cdot}50$ $\chi^2_{11}/11 = 0{\cdot}88$
Partial Adjustment Model	T K $\hat{\sigma}$	102 11 0·541%	T K $\hat{\sigma}$	113 11 0·576%	$\eta_2(11, 91) = 2{\cdot}24$ $\chi^2_{11}/11 = 4{\cdot}11$
Goldfeld Model	T K $\hat{\sigma}$	101 7 0·592%	T K $\hat{\sigma}$	112 7 0·662%	$\eta_2(11, 94) = 3{\cdot}38$ $\chi^2_{11}/11 = 6{\cdot}37$
Risk-adjusted spread model (22)	T K $\hat{\sigma}$	102 12 0·407%	T K $\hat{\sigma}$	113 12 0·391%	$\eta_2(11, 91) = 0{\cdot}32$ $\chi^2_{11}/11 = 0{\cdot}38$

correction mechanism (ECM) by the solution of (22) for the $I(1)$ variables:

$$\mathrm{ECM}_t = m_t - p_t - \tfrac{1}{2}y_t + 5{\cdot}62AS^*_t + 3{\cdot}88AR_{1_t} + 1{\cdot}05(p_t - p_{t-4}) + 1{\cdot}34\Delta\hat{p}_t - 1{\cdot}83R_{nsa_t} - 1{\cdot}41. \quad (24)$$

This represents the $I(0)$ deviations from the long-run money-demand equation, and is close to the residuals from (19). Figure 11 records its behaviour over 1961(1)-1988(3): the ECM shows the largest deviations during "missing money" and "M1-explosion" episodes. Since the ECM is at lag 2, levels variables in (22) are thereby introduced at a 2-period lag, so that the current-dated levels are transformed to two-period differences, and hence to $I(0)$. Re-estimating yields:

$$\begin{aligned}\Delta(m-p)_t = &\underset{(0{\cdot}092)}{-0{\cdot}292}\Delta_4(m-p)_{t-1} \underset{(0{\cdot}024)}{-0{\cdot}238}\mathrm{ECM}_{t-2} \underset{(0{\cdot}091)}{-1.309}\Delta_2^+\mathrm{AS}^*_t \underset{(0{\cdot}043)}{+0.428}\Delta\mathrm{A}y_t \\ &\underset{(0{\cdot}058)}{-1{\cdot}034}\Delta_2^+AR_{1_t} \underset{(0{\cdot}047)}{-0{\cdot}222}\Delta R_{ma_t} \underset{(0{\cdot}296)}{-0{\cdot}842}\Delta_1\Delta_4 p_{t-1} \underset{(0{\cdot}036)}{-0{\cdot}304}\Delta_2^+\Delta\hat{p}_t \underset{(0{\cdot}142)}{+0{\cdot}409}\Delta_2^+R_{nsa_t} \\ &\underset{(0{\cdot}042)}{-0{\cdot}158}\Delta^2(m-p)_{t-4} \underset{(0{\cdot}003)}{+0{\cdot}012}D_t \qquad (25)\end{aligned}$$

$R^2 = 0{\cdot}906, \quad \hat{\sigma} = 0{\cdot}374\%, \quad F(11, 100) = 87{\cdot}7, \quad DW = 1{\cdot}86, \quad SC = -10{\cdot}82.$

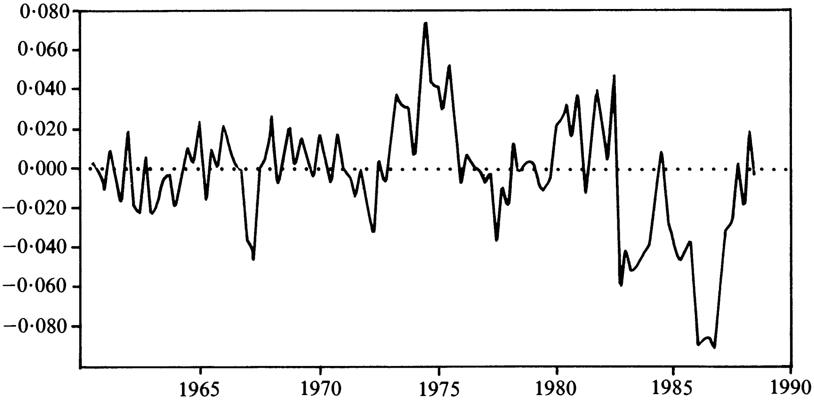

FIGURE 11

Time series of the error correction variable in (24)

In (25), $\Delta_2^+ x_t$ denotes the 2-period difference $(x_t - x_{t-2})$, not divided by 2. Equation (25) is just a reparameterization of (22) to sustain the use of conventional critical values, but it also highlights that real M1 is explained by a combination of short-term dynamic adjustments and the linear function of long-run influences from (24).

VIII. CONCLUSION

Surveying the wreckage after more than a decade of money demand function breakdowns, Goldfeld and Sichel (1990) reluctantly conclude that a constant parameter money demand function may not exist:

> "The repeated breakdown of existing empirical models in the face of newly emerging data has fostered a vast industry devoted to examining and improving the demand for money function."
> "While hardly definitive, these results are certainly suggestive of recurring bouts of instability in money demand."

It is surprising then that the specification presented here is uniformly successful over periods of known difficulty. We conclude that, contrary to the prevailing view, there is good evidence for the existence of a stable, cointegrating money demand function, based on theory, with an error-correction specification.

The present model makes no claim to represent a uniquely correct demand function for M1. It does however, indicate several elements that a money demand function displaying parameter constancy over 1960–1988 may be expected to possess. Four distinctive elements of the present model of money demand appear to be essential in order to track consistently and to accept parameter constancy over the period 1960(3)–1988(3):

(i) Dynamic structure more complex than partial adjustment is required to represent differing adjustment speeds and reactions depending on the source of change. The error-correction model with appropriately lagged independent variables works effectively.

(ii) Representation of the changing complex of yields on monetary instruments is useful. The prevailing yield on interest-bearing instruments in non-transactions M2 is needed in a form that reflects changing availability of instruments and the time required for new instruments to come fully into use by wealth holders. Although our model of adaptation is arbitrary, conditional on maintaining the form and dynamics of (18), Hendry and Ericsson (1991) note estimates of the parameters of the logistic function in (13) (which we set at 7·0 and 0·8) which were 5·5 and 0·65 respectively; the weights are little altered and the other regression coefficients remain essentially the same. This approach could be generalized to allow learning speed to depend on the interest differential. In any event, a model that does not take into account the variation in available yields on monetary instruments deriving from financial innovation will reject parameter constancy, especially for own rates.

(iii) The impact of inflation on money demand should be accounted for. This may partly reflect the dynamics of adjustment of money demand to changes in the price level.

(iv) The impact on money demand of risk and return to long-term bond holding requires representation. Money is held as a safe asset in portfolios in differing amounts corresponding to the return of long-term bond holding and apparent risk from variations in asset values. The increase in the risk-adjusted long-term bond yield appears to be the principal explanation of the Missing Money decline in money demand. The increase in bond risk also appears to be the principal explanation of the Great Velocity Decline. These effects can be summarized in a risk-adjusted long yield. The statistical significance of the bond risk and return variables appears inescapable. In the baseline model the t-statistics on the variables are 13·6, 10·9, and 7·8; in the risk-adjusted yield spread model it is 13·6.

In the absence of such variables, empirical M1 demand models suffer structural breakdown, suggesting alternative measures of money (see e.g. Barnett (1980)), or requiring *ad hoc* rationalizations (e.g. dummy variables) for a downward shift in the amount demanded in 1974–1976 or for upward shifts in 1982–1983 and in 1985–1986. Our specification provides little evidence of a statistically significant shift in the M1 demand function during 1960(3)–1988(3). There is strong evidence that a mis-specified model of M1 demand will suffer a structural breakdown that may be interpreted as a shift when there are substantial changes in the values of variables omitted or in the rapidity of variation in independent variables. In addition to income and short-term interest rate levels, four additional significant aspects of M1 demand are: dynamic structure, maximal own-interest yield on monetary assets, inflation rates, and long-term bond risk and return.

APPENDIX: DATA SOURCES, DATA SERIES AND CREATED VARIABLES

[A] Citibase data tape, data extracted October, 1989.

[B] Federal Reserve Board of Governors.

[C] Salomon Bros. *Analytical History of Yields and Yield Spreads.*

M1, seasonally adjusted, monthly from [A].[16]

R_1, 1-month Treasury bill rate using the discount rate, d, from [C]: monthly beginning of month.[16] Converted to coupon equivalent yield by the formula $R_1 = d/(1-d/1200)$.

R, 20-year Treasury-bond yield to maturity, from [C]: monthly beginning of month.[16]

R_p, Passbook interest rate, quarterly from [B].

R_{cd}, Commercial Bank small CD rate, quarterly, from [B] starting 1965(4).

R_{mf}, Money Market Mutual Fund rate, quarterly from [B] starting 1974(2).

R_{ma}, Learning-adjusted, maximal M2 interest rate, see text.

R_n, NOW account rate, taken to be nil prior to national introduction of NOW's, from [B] starting 1981(1). Previously, NOW accounts were available only in New England (approximately 6% U.S. population and personal income).

16. Conversion of monthly figures to quarterly was by average of monthly figures Jan.–Feb.–March, etc.

R_{sn}, SuperNOW rate, taken to be nil prior to national introduction of SuperNOW's, then from [B] starting 1983(1).

D, Dummy variable for 1980 credit controls: 1980(2): -1 1980(3): $+1$

Y, Real GNP quarterly, price basis 1982, from [A].

P, GNP implicit price deflator quarterly, basis 1982, from [A].

V_m, monthly measure of long-term bond financial risk (volatility), based on H_t, the 1-month holding period yield on 20-year Treasury bonds (at monthly rate) given by:

$$H_t \equiv [D_t R_t/12] - [(D_t - 1)R_{t+1}/12]$$

where

$$D_t = (1 - g_t^{240})/(1 - g_t) \quad \text{and} \quad g_t = [1 + (R_t/1200)]^{-1}.$$

Campbell and Shiller (1983) discuss this linear approximation of holding period yield. The text explains the construction of V from H_t. Difficulties with the approach above are:

(1) there is an anomaly in the 20-year yield data from Salomon Bros. *Analytical History* at January/February 1973, reflecting a change in the available maturities and coupons of Treasury bonds. For the January 1973 holding period yield only, 10 year data were used treating the instrument as a consol.

(2) Salomon Bros. data are not raw yield data but are fitted values to constant maturity of actual yields to maturity.

(3) the formula for holding period yield used is for bonds trading at par whereas actual bond prices vary significantly about par. As a check on our calculation of V_q we used a monthly series of holding period yields on long-term Treasury bonds from the CRSP Bond tape in place of our calculated H_t, to derive V_{crsp}. Inasmuch as they are not generally available, a listing of 20-year bond holding-period yield volatility, V_q, the measure V_a used in BHS, and the alternative V_{crsp} over the BHS sample are available from the authors on request. Results over the BHS sample from substitution of V_{crsp} were essentially identical.

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REFERENCES

ANDO, A. and SHELL, K. (1975), "Demand for Money in a General Portfolio Model in the Presence of an Asset that Dominates Money", in Fromm, G. and Klein, L. (eds.), *The Brookings Model: Prespective and Recent Developments* (New York: North-Holland/American Elsevier), 560-563.

AKERLOF, G. A. (1979), "Irving Fisher on His Head: The Consequences of Constant Target—Threshold Monitoring of Money Holdings", *Quarterly Journal of Economics*, **93**, 169-188.

BABA, Y., HENDRY, D. F. and STARR, R. M. (1985), "US Money Demand, 1960-1984" (Department of Economics Discussion Paper, University of California, San Diego).

BARNETT, W. A. (1980), "Economic Monetary Aggregates: An Application of Index Number and Aggregation Theory", *Journal of Econometrics*, **14**, 11-48.

BAUMOL, W. J. (1952), "The Transactions Demand for Cash: An Inventory Theoretic Approach", *Quarterly Journal of Economics*, **66**, 545-556.

BROWN, R. L., DURBIN, J. and EVANS, J. M. (1975), "Techniques for Testing the Constancy of Regression Relationships over Time", *Journal of the Royal Statistical Society, Series B*, **37**, 149-192 (with discussion).

CAGAN, R. (1956), "The Monetary Dynamics of Hyperinflation", in, M. Friedman, *op. cit.*, 25-57.

CAMPBELL, J. and SHILLER, R. (1983), "A Simple Account of the Behavior of Long-Term Interest Rates" (NBER Working Paper, September).

CAMPOS, J. and ERICSSON, N. R. (1988), "Econometric Modeling of Consumers' Expenditure in Venezuela" (International Finance Discussion Paper No. 325, Federal Reserve Board of Governors, Washington, D.C.).

CHANG, W. W., HAMBERG, D. and HIRATA, J. (1983), "Liquidity Preference as Behavior Toward Risk is a Demand for Short-term Securities—Not Money", *American Economic Review*, **73**, 420-427.

CHONG, Y. Y. and HENDRY, D. F. (1986), "Econometric Evaluation of Linear Macro-Economic Models", *Review of Economic Studies*, **53**, 671-690.

CHOU, Y.-T. (1988), "The Role of Risk in Financial Markets", (Ph.D. dissertation, San Diego: UCSD).

CHOW, G. C. (1960), "Tests of Equality between Sets of Coefficients in Two Linear Regressions", *Econometrica*, **28**, 591-605.

COCHRANE, D. and ORCUTT, G. H. (1949), "Application of Least Squares Regression to Relationships containing Auto-correlated Error Terms", *Journal of the American Statistical Association*, **44**, 32-61.

COOLEY, T. and LEROY, S. (1981), "Identification and Estimation of Money Demand", *American Economic Review*, **71**, 825-844.

CUTHBERTSON, K. (1989), "Modelling the Demand for Money" (Mimeo, Bank of England).

DAVIDSON, J. E. H., HENDRY, D. F., SRBA, F. and YEO, S. (1978), "Econometric Modelling of the Aggregate Time-series Relationship between Consumers' Expenditure and Income in the United Kingdom", *Economic Journal*, **88**, 661-692.

DICKEY, D. A. and FULLER, W. A. (1979), "Distribution of the Estimators for Autoregressive Time Series with a Unit Root", *Journal of the American Statistical Association*, **74**, 427-431.

DICKEY, D. A. and FULLER, W. A. (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", *Econometrica*, **49**, 1057-1072.

DUFOUR, J.-M. (1982), "Recursive Stability Analysis of Linear Regression Relationships: An Exploratory Methodology", *Journal of Econometrics*, **19**, 31-76.

DURBIN, J. and WATSON, G. S. (1950), "Testing for Serial Correlation in Least Squares Regression. I", *Biometrika*, **37**, 409-428.

DURBIN, J. and WATSON, G. S. (1951), "Testing for Serial Correlation in Least Squares Regression. II", *Biometrika*, **38**, 159-178.

ENGLE, R. F. (1982), "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation", *Econometrica*, **50**, 987-1007.

ENGLE, R. F. (1984), "Wald, Likelihood Ratio, and Lagrange Multiplier Tests in Econometrics", Chapter 13 in Griliches, Z. and Intriligator, M. D. (eds.) *Handbook of Econometrics, Volume II* (Amsterdam: North-Holland), 775-826.

ENGLE, R. F. and GRANGER, C. W. J. (1987), "Co-integration and Error Correction: Representation, Estimation, and Testing", *Econometrica*, **55**, 251-276.

ENGLE, R. F. and HENDRY, D. F. (1989), "Testing Super-Exogeneity and Invariance" (Department of Economics Discussion Paper 89-51, University of California, San Diego).

ENGLE, R. F., HENDRY, D. F. and RICHARD, J.-F. (1983), "Exogeneity", *Econometrica*, **51**, 277-304.

EVANS, P. (1984), "The Effects on Output of Money Growth and Interest Rate Volatility in the United States", *Journal of Political Economy*, **92**, 204-222.

FAVERO, C. and HENDRY, D. F. (1989), "Testing the Lucas Critique: A Review" (mimeo, Nuffield College, Oxford).

Federal Reserve Bank of San Francisco (1983) *Monetary Targeting and Velocity Conference Proceedings* (San Francisco: Federal Reserve Bank of San Francisco).

FRIEDMAN, M. (ed.) (1956) *Studies in the Quantity Theory of Money* (Chicago: University of Chicago Press).

FRIEDMAN, M. and SCHWARTZ, A. J. (1982) *Monetary Trends in the United States and the United Kingdom: Their Relation to Income, Prices, and Interest Rates, 1867-1975* (Chicago: University of Chicago Press).

GARCIA, J. and PAK, S. (1979), "Some Clues in the Case of the Missing Money", *American Economic Review*, **69**, 330-334.

GOLDFELD, S. M. (1973), "The Demand for Money Revisited", *Brookings Papers on Economic Activity*, 577-646.

GOLDFELD, S. M. (1976), "The Case of the Missing Money", *Brookings Papers on Economic Activity*, 683-730.

GOLDFELD, S. M. and SICHEL, D. E. (1990), "The Demand for Money", Chapter 8 in Friedman, B. M. and Hahn, F. H. (eds.), *Handbook of Monetary Economics, Volume I* (Amsterdam: North-Holland), 299-356.

GORDON, R. J. (1983), "The 1981-82 Velocity Decline: A Structural Shift in Income or Money Demand?" in Federal Reserve Bank of San Francisco (1983).

GORDON, R. J. (1984), "The Short-run Demand for Money: A Reconsideration", *Journal of Money, Credit and Banking*, **16**, 403-434.

GRILICHES, Z. (1957), "Hybrid Corn: An Exploration in the Economics of Technological Change", *Econometrica*, **25**, 501-522.

HARVEY, A. C. (1981) *The Econometric Analysis of Time Series* (Oxford: Philip Allan).

HENDRY, D. F. (1979), "Predictive Failure and Econometric Modelling in Macroeconomics: The Transactions Demand for Money", Chapter 9 in Ormerod, P. (ed.), *Economic Modelling* (London: Heinemann Education Books), 217-242.

HENDRY, D. F. (1985), "Monetary Economic Myth and Econometric Reality", *Oxford Review of Economic Policy*, **1**, 72-84.

HENDRY, D. F. (ed.) (1986) *Econometric Modelling with Cointegrated Variables, Oxford Bulletin of Economics and Statistics*, **48**, (special issue).

HENDRY, D. F. (1987), "Econometric Methodology: A Personal Perspective", Chapter 10 in Bewley, T. F. (ed.), *Advances in Econometrics, Volume 2* (Cambridge: Cambridge University Press), 29-48.

HENDRY, D. F. (1989), *PC-GIVE: An Interactive Econometric Modelling System, Version 6.0/6.01* (Oxford: University of Oxford, Institute of Economics and Statistics and Nuffield College).

HENDRY, D. F. and ERICSSON, N. R. (1991), "Modeling the Demand for Narrow Money in the United Kingdom and the United States", *European Economic Review*,**35**, 833-881.

HENDRY, D. F. and MIZON, G. E. (1978), "Serial Correlation as a Convenient Simplification, Not a Nuisance: A Comment on a Study of the Demand for Money by the Bank of England", *Economic Journal*, **88**, 549-563.

HENDRY, D. F. and MIZON, G. E. (1990), "Evaluating Dynamic Econometric Models by Encompassing the VAR", forthcoming in Phillips, P. C. B. and Hall, V. B. (eds), *Models, Methods and Applications of Econometrics* (Oxford: Basil Blackwell).

HENDRY, D. F., PAGAN, A. R. and SARGAN, J. D. (1984), "Dynamic Specification", Chapter 18 in Griliches, Z. and Intriligator, M. D. (eds.), *Handbook of Econometrics, Volume II.* (Amsterdam: North-Holland), 1023-1100.

HENDRY, D. F. and RICHARD, J.-F. (1982), "On the Formulation of Empirical Models in Dynamic Econometrics", *Journal of Econometrics*, **20**, 3-33.

HENDRY, D. F. and RICHARD, J.-F. (1983), "The Econometric Analysis of Economic Time Series", *International Statistical Review*, **51**, 111-163 (with discussion).

HENDRY, D. F. and RICHARD, J.-F. (1989), "Recent Developments in the Theory of Encompassing", in Cornet, B. and Tulkens, H. (eds), *Contributions to Operations Research and Economics: the 20th Anniversary of CORE* (Cambridge, Mass.: The MIT Press).

HYLLEBERG, S. and MIZON, G. E. (1989), "Cointegration and Error Correction Mechanisms", *Economic Journal* (Supplement), **99**, 113-125.

JARQUE, C. M. and BERA, A. K. (1980), "Efficient Tests for Normality, Homoscedasticity and Serial Independence of Regression Residuals", *Economics Letters*, **6**, 255-259.

JOHANSEN, S. (1988), "Statistical Analysis of Cointegration Vectors", *Journal of Economic Dynamics and Control*, **12**, 231-254.

JOHANSEN, S. and JUSELIUS, K. (1990), "Maximum Likelihood Estimation and Inference on Cointegration—With Applications to the Demand for Money", *Oxford Bulletin of Economics and Statistics*, **52**, 169-210.

JUDD, J. and MOTLEY, B. (1984), "The 'Great Velocity Decline' of 1982-1983: A Comparative Analysis of M1 and M2", *Federal Reserve Bank of San Francisco Economic Review*, **3**, 56-74.

JUDD, J. P. and SCADDING, J. L. (1982), "The Search for a Stable Money Demand Function: A Survey of the Post-1973 Literature", *Journal of Economic Literature*, **20**, 993-1023.

KIVIET, J. (1987) *Testing Linear Econometric Models* (Amsterdam: University of Amsterdam Press).

KOSKELA, E. and VIREN, M. (1984), "The Goldfeld Demand for Money Equation Revisited" (Bank of Finland Discussion Paper).

LAIDLER, D. (1985) *The Demand for Money: Theories, Evidence, and Problems* (New York: Harper and Row).

MACKINNON, J. G. (1990), "Critical Values for Cointegration Tests" (mimeo, University of California at San Diego).

MILBOURNE, R. (1983), "Optimal Money Holding Under Uncertainty", *International Economic Review*, **24**, 685-698.

MILLER, M. H. and ORR, D. (1966), "A Model of the Demand for Money by Firms", *Quarterly Journal of Economics*, **80**, 735-759.

MCALEER, M., PAGAN, A. R. and VOLKER, P. A. (1985), "What Will Take the Con Out of Econometrics?", *American Economic Review*, **75**, 293-301.

MIZON, G. E. (1984), "The Encompassing Approach in Econometrics", in Hendry, D. F. and Wallis, K. F. (eds.), *Econometrics and Quantitative Economics* (Oxford: Basil Blackwell).

MIZON, G. E. and HENDRY, D. F. (1980), "An Empirical Application and Monte Carlo Analysis of Tests of Dynamic Specification", *Review of Economic Studies*, **47**, 21-45.

MIZON, G. E. and RICHARD, J.-F. (1986), "The Encompassing Principle and Its Application to Testing Non-nested Hypotheses", *Econometrica*, **54**, 657-678.

OFFENBACHER, E. and PORTER, R. (1982), "Update and Extensions on Econometric Properties of Selected Monetary Aggregates" (Board of Governors of the Federal Reserve System, duplicated, Office Correspondence).

OSTERWALD-LENUM, M. (1990), "Recalculated and Extended Tables of the Asymptotic Distribution of some Important Maximum Likelihood Cointegration Test Statistics" (mimeo, University of California at San Diego).

PAGAN, A. R. and ULLAH, A. (1985), "The Econometric Analysis of Models with Risk Terms" (Paper presented at the World Congress of the Econometric Society, Boston).

PHILLIPS, P. C. B. (1991), "Optimal Inference in Cointegrated Systems", *Econometrica*, **59**, 283-306.

PHILLIPS, P. C. B. and LORETAN, M. (1991), "Estimating Long-Run Economic Equilibria", *Review of Economic Studies*, **58**, 407-436.

RAMSEY, J. B. (1969), "Tests for Specification Errors in Classical Linear Least-squares Regression Analysis", *Journal of the Royal Statistical Society, Series B*, **31**, 350-371.

ROSE, A. K. (1985), "An Alternative Approach to the American Demand for Money", *Journal of Money, Credit, and Banking*, **17**, 439-455.

SARGAN, J. D. (1964), "Wages and Prices in the United Kingdom: A Study in Econometric Methodology" in Hart, P. E., Mills, G. and Whitaker, J. K. (eds.), *Econometric Analysis for National Economic Planning* (Colston Papers, Vol. 16) (London: Butterworths), 25-63 (with discussion); reprinted in Hendry, D. F. and Wallis, K. F. (eds.) (1984), *Econometrics and Quantitative Economics* (Oxford: Basil Blackwell), 275-314.

SARGAN, J. D. (1980), "Some Tests of Dynamic Specification for a Single Equation", *Econometrica*, **48**, 878-897.

SIMPSON, T. D. and PARKINSON, P. M. (1984), "Some Implications of Financial Innovations in the United States (Board of Governors of the Federal Reserve System, Staff Paper 139).

SIMPSON, T. D. and PORTER, R. (1980), "Some Issues Involving the Definition and Interpretation of Monetary Aggregates", in *Controling Monetary Aggregates III* (Boston: Federal Reserve Bank of Boston).

SMITH, G. W. (1986), "A Dynamic Baumol-Tobin Model of Money Demand", *Review of Economic Studies*, **53**, 465-469.

SPANOS, A. (1986) *Statistical Foundations of Econometric Modelling* (Cambridge: Cambridge University Press).

STARR, R. (1983), "Variation in the Maturity Composition of Debt and Behavior of the Monetary Aggregates: The Maturity Shift Hypothesis", *Federal Reserve Bank of San Francisco, Economic Review*, Nov 1983, 202-233.

STOCK, J. H. and WATSON, M. W. (1991), "A Simple Estimator of Cointegrating Vectors in Higher-order Integrated Systems" (mimeo, University of California, Berkeley, January).

TATOM, J. A. (1984), "Interest Rate Volatility: Its Link to the Variability of Monetary Growth and Economic Performance", *Federal Reserve Bank of St. Louis Review*, November, 31-34.

TOBIN, J. (1956), "The Interest-elasticity of the Transaction Demand for Cash", *Review of Economics and Statistics*, **38**, 241-247.

TOBIN, J. (1958), "Liquidity Preference as Behavior Toward Risk", *Review of Economic Studies*, **25**, 65-86.

TURK, C. (1984), "Uncertainty and the Demand for Money" (London School of Economics, Centre for Labour Economics, DP 644).

WALSH, C. (1984), "Interest Rate Volatility and Monetary Policy", *Journal of Money Credit and Banking*, **16**, 133-150.

WHITE, H. (1980), "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity", *Econometrica*, **48**, 817-838.

WHITE, H. (1988) *Specification Analysis in Econometrics* (Cambridge: Cambridge University Press) (forthcoming).