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Working Paper Number	1981-012A
Creation Date	January 1981
Citable Link	https://doi.org/10.20955/wp.1981.012
Suggested Citation	Cheng, D.C., Hooks, D.L., 1981; Price Expectations and the Demand for Real Money Balances: Tests of Observed, Adaptive, and Rational Expectations Hypothesis, Federal Reserve Bank of St. Louis Working Paper 1981-012. URL https://doi.org/10.20955/wp.1981.012

PRICE EXPECTATIONS AND THE DEMAND FOR REAL MONEY BALANCES: TESTS OF OBSERVED, ADAPTIVE, AND RATIONAL EXPECTATIONS HYPOTHESIS

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Working Paper Series No. 81-012

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Price Expectations and the Demand for Real Money Balances: Tests of Observed, Adaptive, and Rational Expectations Hypothesis

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I. Introduction

One of the more elusive subjects of macroeconometric investigation in the U.S. since the mid-1970's has been the "missing money" implied by the tendency of conventional aggregate money demand models to overpredict money balance holdings over this period. [1/] To be more precise, much of this research has been concerned with the nature and causes of an apparent increase in the instability of estimated quarterly models since about 1973. [2/] Not only does a lack of a stable relationship between money demand and its presumed determinants--interest rates and income (or wealth)--raise credibility questions concerning a large class of macroeconometric models, but it can pose serious problems for the conduct of monetary policy.

In particular, the existence of frequent, unpredictable shifts in short-run money demand has been cited by critics of a policy of close control over monetary aggregates as a potential source of greater volatility of interest rates. In view of the current experiment in reserve aggregate targeting being conducted by the Federal Reserve System (Fed), and in view of the calls from some quarters for even closer short-run adherence to annual money target paths than has been achieved thus far under the new procedure, the question of the stability of money demand remains an important policy issue.

The purpose of this paper is to investigate the possibility that at least part of the apparent shift (or shifts) in the conventional quarterly demand for money is due to the omission of price expectations as a determinant of desired real money balances. The next section discusses the theoretical role of price expectations as an argument in money demand functions and the results of some preliminary work with survey data and an adaptive expectations model are reported in Section III. Section IV presents a model of rational price expectations and the estimation procedure that are a major focus of this paper, followed by the estimated coefficients, the results of some specification error tests and out-of-sample forecasting tests of alternative specifications of money demand functions. The concluding section reviews the findings of this study and discusses their implications.

II. The Role of Price Expectations

Most of the empirical models in the money demand literature are of the general form

(1) $$m = \alpha_0 + \alpha_1 y + \alpha_2 i,$$

where m is real money balances, y is real income or wealth, and i is the nominal return on other financial assets (bonds). To the extent price expectations are considered at all in these studies, they usually are explicitly or implicitly assumed to be fully reflected in observed nominal interest rates via the "Fisher Effect". 3/ However, it is possible for price expectations to enter as an argument in equation (1) if: (a) there is less than a full Fisher Effect, (b) real goods are substitutes for money balances, or (c) desired real money balances are determined by real returns on financial assets. Cases (a) and (b) suggest a price expectations variable π should be added to equation (1), giving

(2) $$m = \alpha_0 + \alpha_1 y + \alpha_2 i + \alpha_3 \pi,$$

where $\alpha_3 < 0$. 4/

If, on the other hand, we follow Tobin [1969] and assume money demand is a function of its own real return (r_m) and real alternative financial asset rates (r), then (2) can be written

(3) $$m = \alpha_0 + \alpha_1 y + \alpha_2 (i - \pi) + \alpha_3(i_m - \pi)$$

or

(4) $$m = \alpha_0 + \alpha_1 y + \alpha_2 r + \alpha_3 r_m ,$$

where i_m is the nominal return on money balances. Equation (3) is obtained by assuming:

$$1 + i = (1 + r)(1 + \pi) = 1 + r + \pi + r\pi,$$
$$i = r + \pi + r\pi$$
$$= r(1 + \pi) + \pi,$$

$$\therefore r = (i - \pi)/(1 + \pi).$$

If $r\pi$ is sufficiently small, $r = (i-\pi)$ and we have the terms in (3) and (4). Of course, the greater is π, the greater the error in omitting this product. 5/ If i_m is not an explicit return or if we follow Dutton [1979] and derive a function with a real rate <u>and</u> price expectations as arguments, we can rewrite (3) as

(5) $m = \alpha_0 + \alpha_1 y + \alpha_2 (i - \pi) + \alpha_3 \pi$.

Note that rearranging terms in (5) gives

(6) $m = \alpha_0 + \alpha_1 y + \alpha_2 i + (\alpha_3 - \alpha_2) \pi$,

where the expected sign of the coefficient of π depends upon the relative magnitudes of α_3 and α_2. If $\frac{\partial m}{\partial r} = \alpha_2 > \frac{\partial m}{\partial \pi} = \alpha_3$, then $(\alpha_3 - \alpha_2) < 0$. As Albon and Valentine [1978] pointed out, however, although the theoretical assumptions underlying equations (2) and (6) are different, coefficient estimates will not depend upon which theory is invoked; thus, they cannot be used to allow the researcher to choose between the competing hypotheses. 6/

If (1) is estimated when (2) is the true model the omission of π results in a specification error that can bias the estimated coefficients of (1). The effects of this error can be determined by regressing the omitted variable on the included explanatory variables of equation (1)

(7) $\pi = \beta_0 + \beta_1 y + \beta_2 i + \mu$,

and noting that $E(\hat{\alpha}_2) = \alpha_2 + (\alpha_3 \beta_2)$.

If $\beta_2 > 0$ (<0) and $\alpha_3 < 0$ (>0), then $(\alpha_3 \beta_2) < 0$ (>0). Since $\partial m/\partial i$ and $\partial m/\partial \pi$ are assumed to be negative, omission of π will result in either a negative or a positive bias in the estimate of the coefficient of the nominal interest rate.

Another specification error could result if the true opportunity cost is the real rate $r(= i - \pi)$, but $i(= r + \pi)$ is used instead. In this case, $E(\hat{\alpha}_2) = \alpha_2 - (\alpha_2 \beta_2)$ and, if $\alpha_2 < 0$ and $\beta_2 > 0$ (<), estimation of equation (1) would result in a positive (negative) bias in the estimated interest rate coefficient. It is possible that biased coefficient estimates due to specification errors such as these could have been responsible in part for Goldfeld's finding (which he did not report) that a long-term rate was either not significant or did not have the expected sign, especially if long rates are more sensitive to price expectations than are short rates. Some tests for the effects of possible specification errors are conducted in a later section of this paper.

III. Empirical Results with Observed and Adaptive Price Expectations.

A. Model Specification

This study follows convention in taking the work of Goldfeld [1973] as a point of departure. His basic model specified desired real money balances as a function of real income and interest rates

$$(8)\ \ln m_t^* = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 \ln i_t + \epsilon_t,$$

where ln denotes natural logarithms. But, since m_t^* is not observable, it was assumed that individuals only partially adjust actual to desired money balances within a quarter due to adjustment costs; thus,

$$(9)\ \ln m_t - \ln m_{t-1} = \lambda(\ln m_t^* - \ln m_{t-1}),\ 0 < \lambda < 1$$

where λ is the speed of adjustment.

Solving (9) for $\ln m_t^*$ and substituting into 8 gives

$$(10)\ \ln m_t = \lambda\,\alpha_0 + \lambda\,\alpha_1 \ln y_t + \lambda\,\alpha_2 \ln i_t + (1-\lambda)\ln m_{t-1} + \lambda\,\epsilon_t.$$

The version preferred by Goldfeld in most estimates was

$$(11)\ \ln m_t = \alpha_0 + \alpha_1 \ln \frac{Y_t}{P_t} + \alpha_2 \ln CPR_t + \alpha_3 \ln TD_t + \alpha_4 \ln \frac{M_{t-1}}{P_{t-1}} + u_t ,$$

where CPR is the commercial paper rate and TD is a time deposit rate. 7/ Goldfeld included these two short-term rates on the assumption that holders of money balances do not all face the same opportunity costs. Income is Gross National Product (GNP) and M is the narrow M1 money stock, both deflated by the 1972-based GNP deflator, P. The M1 measure is used in order to avoid the complications caused by a measure in which some components do have an explicit yield and to abstract from the current issue of the proper measure of the money stock during a period of rapid financial innovation.

Estimates of this model are presented in Table 1. Models 1.1 - 1.4 are reproduced from Goldfeld [1973]. He broke his sample period estimates at 1961.4 in Models 1.2 and 1.3 in order to test for the possible effects of the introduction of large certificates of deposit on the stability of the demand for transactions balances; the hypothesis of instability was rejected. As one can see, however, the time deposit rate is not significant in the later subperiod. Model 1.4 is representative of Goldfeld's attempt to introduce price expectations into his model. Even though it was significant and had the expected sign, this variable, which was either the actual last quarter percent change in the price index, or a proxy constructed from survey data, was not included in subsequent specifications for reasons to be discussed later in this paper.

Model 1.5, from Goldfeld [1976], is the one whose poor out-of-sample forecasting performance, in contrast to his findings in the 1973 paper, prompted Goldfeld and others to search for the missing money. This model produced a large static root-mean-squared forecasting error (RMSE) over ten quarters beginning in 1974.1 of 4.29 (Goldfeld [1976, p. 686]), and it overpredicted desired money balances in each quarter.

The remaining models in Table 1 are replications and extensions of Goldfeld's work from an earlier investigation into the role of price expectations in money demand, which produced reasonably close approximations to his reported estimates; moreover, they confirmed his unreported finding that the long term interest rate did not perform well over his original estimation period, 1952.2 to 1972.4. [8/] Some previous work on the interrelationships among interest rates, the money stock, and the price level prompted us to reestimate this model over two subperiods with mid-1965 as the break point. [9/]

The supposition that the erroneous use of nominal interest rates would have a greater effect on coefficient estimates for long-term bond rates during periods of changing inflationary expectations appear to be borne out in the estimates of Models 1.6 - 1.8. Only in the earlier period of relatively low actual and, presumably, expected inflation does this coefficient exhibit a significant negative value. Apparently, the specification error bias that is readily observed in the second period (Model 1.8) dominates the estimates for the entire Goldfeld period. Another problem appears in the estimate of Model 1.11, in which the short-term rate is reintroduced. Neither the CPR nor the TD rates are significant in the post-1965 period, suggesting a multicollinearity problem with a pattern that changes over time.

This possibility has been explored in greater detail elsewhere (Cheng and Hooks [1979]) with principal components regression (PCR). The time deposit rate itself was found to be the source of the multicollinearity problem; indeed, PCR estimates of this coefficient did not have a significant negative value in either of the two subperiods with or without the inclusion of the long-term rate. If the pattern of multicollinearity among the money demand determinants is changing over time, the out-of-sample forecasting error can be increased in addition to the effect of collinearity on the coefficient estimates themselves. [10/] Because the focus of this study is on the significance and sign of the interest rate and price expectations coefficients as well as on relative model forecasting performance, it was decided to eliminate the TD rate in order to avoid the potential biases and errors due to multicollinearity. Most of the estimates in the tables that follow are for models including only CPR.

Another change in the specification of the basic Goldfeld model pertains to functional form. Most of our exploratory work on the role of nominal and real interest rates and expected inflation rates was conducted using a linear model, which appears to lend itself more readily to the evaluation of the potential specification errors due to omitted variables discussed above. In addition, since our proxies for real rates and expected price changes could take on negative values, this would necessitate our using a semi-log model. There appears to be no a priori case for either specification in theory, but the reader can judge for himself by comparing the estimates of a linear model (1.12 and 1.13) with those of the log-level form in Table 1 in terms of explanatory power and goodness of fit. 11/

B. Price Expectations from Survey Data

Table 2 presents some representative estimates using survey expectations data and real rate proxies. The two expectations proxies used were from the University of Michigan Survey Research Center and differed in the way in which the response was given by those interviewed. One proxy consisted of responses in terms of ranges of inflation rates; the other consisted of specific values for expected rates of price level change. The results are very similar using both measures. The price expectations variable has a significant negative coefficient in every model with the exception of Model 2.6, which was estimated over the post-1965 inflationary period. Note in particular that there is very little evidence that the inclusion of this expectations variable had the hypothesized effect on the estimated long-term bond coefficient when Models 2.4 - 2.6 are compared with Models 1.6 - 1.8, although the change in Table 2 is in the expected direction. To the extent these proxies capture price expectations, it appears these expectations affect desired real balances through a channel in addition to the Fisher effect on nominal interest rates.

C. Real After-Tax Interest Rates

Several regressions were also run in an attempt to test the hypothesis that the real after-tax rate is the opportunity cost of money balance holdings. The tax effect on interest rates can be evaluated by deriving the real after-tax rate

$$(12) \qquad r^{*} = (1 - \tau)(i - \pi) \qquad 0 \lesseqgtr t \lesseqgtr 1$$
$$= (1 - \tau)\, r,$$

where τ is the marginal tax rate. This means the use of a nominal rate in estimates of equation (1) implies the omission of $\pi/(1-\tau)$ if the true model is

$$(13) \quad m = \alpha_0 + \alpha_1 y + \alpha_2 \{i - [\pi/(1-\tau)]^{\iota}$$
$$= \alpha_0 + \alpha_1 y + \alpha_2 i - \alpha_2 [\pi/(1-\tau)].$$

Equation (7) then becomes

$$(7') \quad \pi/(1-\tau) = \beta_0 + \beta_1 y + \beta_2 i + n,$$

which adds an additional bias to estimates of equation (1), depending on the magnitude of τ. Darby [1976] has estimated the value of τ to range from .25 to .4 for the U.S. 12/ Following Boskin [1978], we used a tax-exempt municipal bond rate as a proxy for the real after-tax rate. 13/ These results were not encouraging, however, and are not reported here.

D. Adaptive Price Expectations

There are several reasons why household survey data may not provide good expectations proxies for aggregate money demand estimations, especially during periods in which actual inflation rates are high and variable by U.S. postwar standards. Hafer and Resler [1980], for example, found that one set of survey forecasts is not rational in that available information apparently was not used by respondents; moreover, it appears survey expectations adapt slowly.

As an alternative, we experimented with an expectations model that involves a straight-forward interpretation of Fisher; i.e., we assumed the observed nominal interest rate can be expressed as

(14) $i_t = \alpha + \beta \pi_t + u_t$,

where $\alpha + u_t$ represents the real rate and $\beta=1$, which implies a full Fisher effect (ignoring taxes). But, if the expected inflation rate is formed according to an adaptive expectations mechanism, we have

$$\pi_t - \pi_{t-1} = \gamma(\dot{P}_t - \pi_{t-1}),$$

$$\therefore \quad \pi_t - \pi_{t-1} + \gamma \pi_{t-1} = \gamma \dot{P}_t$$

(15)

$$\pi_t - (1-\gamma)\pi_{t-1} = \gamma \dot{P}_t$$

$$[1 - (1 - \gamma) D] \pi_t = \gamma \dot{P}_t ,$$

where $\dot{P}_t$ is the actual inflation rate and D is the lag operator. Solving for π_t

(16) $\pi_t = \frac{\gamma}{1 - (1 - \gamma) D} \dot{P}_t$, and substitution of equation

(16) into (14) gives

(17) $i_t = \alpha + \beta \frac{\gamma}{1 - (1 - \gamma) D} \dot{P}_t + u_t$.

In order to obtain an estimate of γ, we write

(18) $i_t - (1-\gamma)\, i_{t-1} = \alpha[1 - (1 - \gamma)D] + \beta \gamma \dot{P}_{t-1} + u_t [1 - (1 - \gamma) D],$

$\therefore\ i_t = \gamma \alpha + (1 - \gamma)\, i_{t-1} + \beta \gamma \dot{P}_t + u_t - (1 - \gamma)\, u_{t-1}$.

Thus $\hat{\gamma}$ is equal to one minus the estimated coefficient of i_{t-1}. Following Feldstein [1970, p. 1335], π_t becomes

(19) $\pi_t = \frac{\gamma}{1-(1-\gamma)^T} \sum_{j=0}^{T} (1-\gamma)^j \dot{P}_{t-j}$,

which we then used in estimating the models reported in Table 3.

Again the results were mixed regarding the hypotheses discussed in Section II. The estimates of the interest rate coefficient are still sensitive to the sample period chosen and, now, to the length of the lag used in obtaining the expectations proxy π. In Models 3.1-3.4, T was set equal to 20, and in 3.5 and 3.6 it was set at 10 quarters. These models were also subjected to out-of-sample forecasting tests despite the limited improvement in the estimates, and the RMSE for the static forecasts of levels of money balances over 12 quarters beginning in 1974.1 are reported in the last column. These can be compared with an RMSE of 6.6 for our estimate of Goldfeld's specification over the same period.

Proxies for inflationary expectations based on weighted averages of past inflation rates are suspect during periods of increasing actual inflation rates which may explain the limited improvement reported here.

In particular, if the actual rate (P) has an upward time trend, the expected rate proxy (π) must still be lower than the observed rate in the period in which expectations are formed in the adaptive model above. Indeed, it would seem to be inconsistent with rational expectations for individuals to continue to form expectations according to such a mechanical rule when additional information about the price trend is available. 14/ Thus, we turn to the main focus of this paper, the estimation of the demand for real balances incorporating rational price expectations.

IV. Rational Price Expectations: Estimates and Tests

A. The Model and Estimation Procedure

Following the work of Muth [1961], we assume that a rational expectation in period t of the inflation rate π in period t+1, conditional on all available informaion Φ in period t, can be modeled

$$(20)\quad \pi_{t+1} = E_t\,(\pi_{t+1} \mid \phi_t) = \pi_{t+1} + v_{t+1},$$

which states that the realized π_{t+1} is equal to the mathematical expectation of future inflation plus an error term with a zero mean that is independent of ϕ. If perfect foresight is assumed, $v_{t+1} = 0$, and one can use the actual inflation rate as π_{t+1}. If not, since the realized π_{t+1} is distributed around the mathematical expectation, then the use of actual inflation rates amounts to an error-in-variables problem that can result in inconsistent parameter estimates. 15/ Therefore, in addition to using actual values of π, we follow McCallum [1976] in using an instrumental variable approach to first estimate a predicted value of π_{t+1} and then including $\hat{\pi}_{t+1}$ as a variable in the money demand function. 16/

The structure of the economy and the nature of national economic policies determine in part what constitutes relevant economic information regarding future rates of price change and, therefore, what variables one should include as instruments in the first stage regression. Information costs may also be important since the rationality of a decision rule depends upon these costs. For example, McCallum terms a rule in which only lagged inflation rates are included in ϕ_t as "partly rational" expectations.

The quarterly price change at annual rates in quarter t+1 is defined as

$$(21) \quad \pi_{t+1} \equiv \left[\left(\frac{P_{t+1}}{P_t} \right)^4 - 1 \right] X\ 100,$$

where P is the GNP deflator. 17/ A number of first stage regression models were tested in order to obtain $\hat{\pi}_{t+1}$, and the one used in the second stage results reported below had the following instrumental variables

$$(22) \quad \hat{\pi}_{t+1} = f\ (y_t,\ CPR_t,\ m_{t-1},\ M_t,\ M_{t-1},\ G,\ \pi_t,\ \pi_{t-1}$$
$$c_{t-1},\ t,\ (^{y}/_{q})_t)\ ,$$

where the first three variables are from the second stage model to be estimated and are defined above, M is nominal money, G is nominal government spending, c is real consumption, t is a time trend, and $(^y/_q)$ is the ratio of real GNP to potential output.[18/]

B. Results

Second stage estimates are presented in Table 4, along with our reestimation of the original Goldfeld specification (4.1) and a benchmark model without price expectations or a time deposit rate (4.2) for purposes of comparison. Both the predicted and the actual current inflation rates are significant in the level specification estimates (4.3 and 4.5, respectively), but the next period expectation is not. An F-test indicates that the addition of $\hat{\pi}$ significantly increases the explanatory power of the model. Note, too, that the use of the actual inflation rate appears to result in a bias in the estimate of the coefficient of π due to the measurement error discussed in the previous section.

Estimation of 4.3 in logarithmic form results in a slightly better fit, as can be seen by examining 4.7. Both predicted current and future inflation rates are significant in models 4.8 and 4.9, which are estimates of a first-difference-in-logarithms specification, although the next period expectations produces a better fit to the data. The first difference or delta-log specification has been suggested by Hafer and Hein [1980] because the Cochrane-Orcutt procedure will underestimate the serial correlation coefficient of the residuals when a lagged dependent variable is a regressor; moreover, they found that estimates of this specification were more stable over time than were the log-level estimates of the conventional money demand function.[19/]

An economic interpretation of the first-difference specification can also be offered. If the rate at which individuals adjust their real money balances is inversely related to changes in the expected rate of inflaton, then $\Delta\pi$ will have a negative coefficient, as is the case in the estimates reported here.

C. Specification Error Tests

An auxillary regression was run in order to evaluate the potential effects of omitting the price expectations variable or misspecifying the opportunity costs of holding real money balances. The auxillary regression for the tests was

$$(23)\quad \pi = \underset{(9.47)}{-21.74} - \underset{(.59)}{5.27}\, y + \underset{(4.74)}{.355}\, CPR + \underset{(8.67)}{.108}\, m_{-1} \qquad \bar{R}^2 = .845 \quad F = 154.9$$

If the error is solely due to the use of a nominal rather than a real rate of interest, then α_2 in equation (6) should be equal to $-(\alpha_3 - \alpha_2)$. This hypothesis was rejected in tests for the equivalence of the estimated coefficient on CPR and $\hat{\pi}$ in models 4.3 and 4.4. However, if the specification error is due to the omission of π as a separate explanatory variable when the true model is (2), the discussion of auxillary regression (7) in conjunction with knowledge of β_2 in (23) indicates that the included nominal rate will have a negative bias. Recall that it is not possible to distinguish between the hypotheses underlying equations (2) and (4). It might also be noted that since the lagged dependent variable is expected to have a positive coefficient and since its coefficient estimate in auxillary regression (23) is also positive, omitting π will result in that variable's coefficient having a negative bias in the case of the first error, and a positive bias in the second.

D. Nominal Adjustment Model

Thus far all of the models reported and evaluated in this paper have assumed a real adjustment specification. Recently, White [1978] has presented an argument that the true specification of the adjustment mechanism is a nominal adjustment process in which the lagged dependent variable should be deflated by the current price level, rather than the lagged price index. 20/ Thus equation (9) is rewritten as

$$(9') \ \ln m_t - \ln m_{t-1} = \lambda\,(\ln m_t^* - \ln m_{t-1}) - (1-\lambda)\,\Delta \ln P_t,$$

and (10) becomes

$$(10') \ \ln m_t = \lambda\,\alpha_0 + \lambda\alpha_1 \ln y_t + \lambda\alpha_2 \ln i_t + (1-\lambda) \ln \frac{M_{t-1}}{P_t} + u_t.$$

White [1978, p. 599] then argued that adding the current inflation rate ($\Delta \ln P_t$) to the real adjustment model as a proxy for price expectations (or any proxy positively correlated with that rate) would result in it having a significant negative coefficient because it simply serves to correct the specification error and not necassarily because the variable plays any causal role in money demand. Adding $\Delta \ln P_t$ to the "correct" nominal adjustment specification, however, should result in a nonsignificant coefficient, which Goldfeld [1973] found in his brief experiment with price expectations. This is because in adding the expectations term to (10'), one is essentially estimating (in our notation) 21/

$$(10'') \ \ln m_t = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 \ln i_t + \alpha_3 \ln m_{t-1}$$
$$+ \alpha_4\,\Delta \ln P_t - \alpha_5\,\Delta \ln P_t + u_t\ .$$

We estimated nominal adjustment versions of our model 4.7 in Table 4 using predicted current and future inflation, respectively, in order to test White's arguments:

$$(24)\quad -.471 + \underset{(3.25)}{} \; .0402\, y_t - .013\, CPR_t + 1.043\, \frac{M_{t-1}}{P_t} - .001\, \hat{\pi}_t$$
$$(3.25)\quad (5.89)\quad (3.89)\quad (33.35)\quad (1.402)$$

$$(25)\quad -.652 + .041\, y_t - .013\, CPR_t + 1.076\, \frac{M_{t-1}}{P_t} - .002\, \hat{\pi}_{t+1}$$
$$(3.38)\quad (6.13)\quad (4.26)\quad (27.41)\quad (2.03)$$

Although there is no clear evidence in these estimates for White's contention that the use of a real adjustment specification as in Table 4 results in biased estimates of the speed of adjustment coefficient λ, there is some support for his argument regarding the role of the current inflation rate as an expectations proxy. Both the magnitude and the singificance of the coefficient of $\hat{\pi}_t$ are less in (24) than in (5.7) in Table 5. On the other hand, the coefficient on $\hat{\pi}_{t+1}$ in (25) is significant at the 5 percent level. 22/ Since the question of the correct specification of the adjustment process is not easily resolved, both real and nominal adjustment models will be subjected to the static forecast tests in the following section.

E. Forecasting Error Tests

Although the partial adjustment model used in this study can be viewed as a dynamic model, and dynamic forecasting errors are often tested in the literature on money demand stability, static forecasting performance results are reported here for several reasons. First, as Hein [1980] has shown, both dynamic and static forecasts are unbiased, but the former are inefficient relative to the latter and this inefficiency increases with the length of the forecast period. Second, it has also been shown (Hein [1980]) that the dynamic forecast error for any period can be obtained from the static forecast errors and the coefficient on the lagged dependent variable.

Results of forecasting tests of the estimated models in Table 4 are presented in Table 5. Although all of the models tested outperform the Goldfeld specificaion (5.1) in terms of RMSE for levels of money balances, fraction of error due to bias, and the ratio of RMSE to the standard error of the estimate (which allows for comparisons across models with different functional form), the best real adjustment model is clearly 5.7. This model is the log-linear form of 4.3, which includes the predicted current inflation rate proxy for rational price expectations. This model's predictions track actual desired money balances quite well over the post-sample period with no systematic over- or underprediction. In comparison, the Goldfeld specification still consistently over-predicts, as does the benchmark model.

Forecasting tests were also performed on (24) and (25) and summary results are reported in rows 5.10 and 5.11, respectively, in Table 5. The nominal adjustment model including the predicted current inflation rate performs only slightly better in terms of the criteria used in these tests than the real adjustment model (4.7 and 5.7), despite having a lower t-statistic for π_t than in the former model.

V. Summary and Conclusions

This paper has reported the results of a study of the role of alternative specifications of a price expectations variable in the demand for real money balances. The addition of a proxy for rational inflation rate forecasts was found to improve both the explanatory power and out-of-sample forecasting performance of a conventional quarterly money demand function. Specification error tests indicated that the expectations variable entered primarily as a separate explanatory variable rather than as an incomplete Fisher Effect on the nominal interest rate. As such, its omission results in a negative bias in the estimated nominal rate coefficient, which could explain the consistent overprediction of desired real balance holdings by conventional models.

Some evidence was also found in support of the specification of a nominal rather than the usual real adjustment process; however, the log-level form of both specifications of the adjustment process produced better fits to the data and better forecasts than the model estimated in level or in first difference forms. Other specification questions aside, however, rational price expectations appears to be an important omitted variable in previous estimates of quarterly money demand functions for the U.S. Not only does the addition of this variable explain some of the previously reported "missing money", but it eliminates some of the apparent instability of the conventional model that has raised questions regarding a policy of close short-run money stock control.

FOOTNOTES

* The research assistance of James Cooley is gratefully acknowledged.

1/ Goldfeld [1976], Enzler, Johnson and Paulus [1976], and Hamburger [1977] were among the first to discover the "missing money" and to undertake a search for the sources of the increased instability of the conventional models.

2/ Hafer and Hein [1980] have presented evidence that at least part of the apparent instability is due to the functional form and estimating technique used in many of the studies; hence, a properly specified model exhibits greater stability. We will consider this issue in a later section.

3/ Heller and Khan [1979, p. 111], for example, made this explicit assumption due to "historically moderate rates of inflation in the U.S."

4/ See Friedman [1956] and Turnovsky [1974] and [1977] for models in this spirit. Although Goldfeld [1973] experimented with this specification, he did not elaborate on its theoretical foundation and subsequently abandoned it for reasons to be discussed later.

5/ Note that if taxes are considered and τ is the marginal tax rate on interest income $(o \leq \tau \leq 1)$, the real after-tax opportunity cost of holding money is $r^* = (1-\tau)(1-\pi)$. Some preliminary work with this variable is dissussed in Section III.

6/ However, estimation of (2) and (5) may produce different results if π and i are collinear. Albon and Valentine [1978] estimated specification (3) and Valentine [1977] estimated (2), both using Australian data. Because this study uses the narrow M1 measure of transactions balances for the U.S., which did not have an explicit return, we did not attempt to estimate (3).

7/ The time deposit rate series was kindly provided by Stephen Goldfeld.

8/ Some of the results in this section were reported in papers presented at the 1979 meetings of the Western and Southern Economic Associations, and at the Federal Reserve Bank of St. Louis.

9/ This choice of periods was suggested by Tobin [1974], who reported an increased sensitivity of equity prices to interest rates since 1965, presumably due to an increased awareness of monetary policy by investors; moreover, a marked increase in the level and variation of the U.S. postwar inflation rate (actual and expected) corresponds roughly to the post-1965 period. Evidence of structural changes in the relationships among money growth, the price level, and financial asset yields and prices are reported in Hooks and Cheng [1978].

10/ Watson and White [1976] attributed this possibility to a changing term structure and attempted to obtain more stable estimates by using ridge regression. Heller and Khan [1979], on the other hand, attacked the problem directly by attempting to model the term structure of rates itself.

11/ Albon and Valentine [1978] used a linear model for the same reasons. Later in this paper we return to the question of functional form in evaluating relative out-of-sample forecasting performance of alternative models.

12/ Of course, if price expectations and marginal tax rates differ among market participants, more complicated expressions will result. On this see Gandolfi [1976].

13/ A time series regression of a municipal bond rate of comparable maturity on the AAA corporate bond rate yielded an estimated coefficient of .72, which implies $\tau \simeq .3$ on average over the sample period.

14/ See Vane and Thompson [1979, pp. 104-105] for a discussion of this point. Of course it is possible that information costs could be so high as to make adaptive rules economically rational.

15/ Use of actual π_{t+1} rather than (20) in (8) is tantamount to introducing v into the money demand error term, ε, which means ε' is not independent of the explanatory variables.

16/ McCallum also noted that π_t (or even π_{t+2} might be appropriate for the purpose of modeling rational expectations. We did experiment with $\hat{\pi}_t$

on the assumption that not even the current quarterly inflation rate, using published price indices, could be known to individuals during the quarter in which they are making decisions regarding desired money balances and how much adjustement to make. Use of the contemporaneous actual inflation rate is also consistent with the argument that information costs can make current or even past rates of change in price economically relevant and, therefore, rational information.

17/ For contemporaneous inflation equation (21) is

$$\pi_t \equiv \left[\left(\frac{P_t}{P_{t-1}} \right)^4 - 1 \right] X\ 100$$

18/ These and the other instruments that were tested were suggested by McCallum's work. They can be obtained by assuming the money demand function is part of a macroeconomic system in which price level changes are endogenous, as are y_t and CPR_t. Of course, when $\hat{\pi}_t$ is estimated, π_t is omitted from the right hand side of equation (22). Log-likelihood tests for stability were performed on (22) (see Quandt [1960]) and there was a weak indication of a change in structure in 1971. The inclusion of dummy time period variables in the first stage regression did not affect the results.

19/ Underestimation of the value of the serial correlation coefficient will result in coefficient estimates that are inconsistent and inefficient (see Theil [1971]).

20/ White [1978, p. 568] attributes the specification error to a disregarding of any "discrepancy between desired and actual money holdings caused by a change in the price level between period t-1 and period t that altered the real value in t of the stock of money already held in t-1".

21/ The potential specification error can be seen by substituting (1) into (9) in log form and adding an error term

$$\ln m_t - \ln m_{t-1} = \lambda[\alpha_0 + \alpha_1 \ln i_t + \alpha_2 \ln y_t + \varepsilon_t) - \ln m_{t-1}] - (1-\lambda)\ \Delta \ln P_t,$$

$$\ln m_t = \lambda\ \alpha_0 + \lambda\ \alpha_1 \ln i_t + \lambda\ \alpha_2 \ln y_t + \lambda\ \varepsilon_t - \lambda \ln m_{t-1} - (1-\lambda)\ \Delta \ln P_t + \ln m_{t-1}$$

$$= \lambda\ \alpha_0 + \lambda\ \alpha_1 \ln i_t + \lambda\ \alpha_2 \ln y_t + (1-\lambda)\ [\ln m_{t-1} - (\ln P_t - \ln P_{t-1})] + \lambda\ \varepsilon_t$$

$$= \lambda\ \alpha_0 + \lambda\ \alpha_1 \ln i_t + \lambda\ \alpha_2 \ln_{yt} + (1-\lambda)\ [\ln M_{t-1} - \ln P_t] + \lambda\ \varepsilon_t\ ,$$

or equation (10') in the text.

22/ If individuals do not know the actual inflation rate during the quarter in which the adjustment is made, the potential bias due to this error may be reduced.

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Table 1

Estimates of Goldfeld's Money Demand Specification

Model/Period	Coefficients: Real Income	Lagged Money	Commercial Paper Rate	Time Deposit Rate	AAA Rate	Price Change	$\bar{R}^2$	F	Standard Error
1.1 1952.2-1972.4	.193 (5.3)	.717 (11.5)	-.019 (6.0)	-.045 (6.01)			.995		.0043
1.2 1952.2-1961.4	.216 (4.6)	.604 (6.4)	-.019 (5.4)	-.060 (4.1)			.978		.0043
1.3 1962.1-1972.4	.191 (3.3)	.632 (4.8)	-.014 (2.4)	-.010 (.3)			.992		.0050
1.4 1952.2-1972.4	.166 (4.9)	.782 (13.1)	-.015 (5.0)	-.038 (3.6)		-6.57 (4.2)	.996		.0042
1.5 1952.2-1973.4	.179 (5.4)	.676 (10.0)	-.018 (6.5)	-.042 (4.0)			.995		.0042
1.6 1952.2-1972.4	.204 (4.8)	.613 (7.2)		-.051 (3.2)	-.018 (1.5)		.995	2115.2	
1.7 1952.2-1965.2	.223 (4.13)	.516 (4.66)		-.045 (3.01)	-.056 (3.14)		.991	272.2	
1.8 1965.3-1972.4	.302 (3.46)	.436 (2.51)		-.028 (.556)	-.0012 (0.63)		.967	169.9	

Table 1 (continued)

	Coefficients							
Model/Period	Real Income	Lagged Money	Commercial Paper Rate	Time Deposit Rate	AAA Rate	Price Change	$\bar{R}^2$	F
1.9 1952.2-1973.4	.179 (5.8)	.678 (9.6)	-.017 (5.6)	-.043 (3.9)			.993	2611.9
1.10 1952.2-1965.2	.189 (4.5)	.705 (9.2)	-.019 (6.2)	-.044 (3.52)			.968	352.4
1.11 1965.3-1973.4	.282 (3.2)	.473 (2.7)	-.005 (.66)	-.031 (.77)			.960	172.3
1.12 1952.2-1972.4	.034 (4.13)	.773 (11.8)	-.770 (4.2)	-2.434 (2.8)			.992	2293.3
1.13 1952.2-1973.4	.028 (3.4)	.798 (11.8)	-.765 (4.4)	-1.821 (2.1)			.997	2555.5

Notes: The constants are not reported. Values in parentheses are t-statistics. All variables are in natural logs except in models 1-12 and 1-13.

Table 2

Money Demand Estimates With Observed Price Expectations

Model / Period	Real Income	Lagged Money	Commercial Paper Rate	AAA Rate	Price Expectations	$\bar{R}^2$	F
			Coefficients				
2.1	.018	.891	-.583		-1.505	.992	2599.8
1952.2-1973.4	(4.8)	(23.7)	(3.2)		(4.1)		
2.2	.018	.926	-1.084		-1.85	.961	286.5
1952.2-1965.2	(3.9)	(22.6)	(3.9)		(3.5)		
2.3	.027	.686	-.396		-1.527	.969	217.8
1965.3-1973.4	(1.8)	(4.2)	(1.5)		(2.8)		
2.4	.019	.847		-.213	-1.567	.992	2374.4
1952.2-1973.4	(3.9)	(17.9)		(.50)	(2.9)		
2.5	.018	.931		-1.114	-1.669	.961	283.5
1952.2-1965.2	(3.3)	(22.4)		(3.8)	(3.5)		
2.6	.027	.622		.296	-.932	.968	204.4
1965.3-1973.4	(1.7)	(3.6)		(.41)	(1.1)		

Note: The constants are not reported. Values in parentheses are t-statistics. All models are estimated in level form. The price expectations proxy is the University of Michigan Survey Research Center measure of price expectations.

Table 3

Money Demand Estimates With Adaptive Price Expectations

Model / Period	Real Income	Lagged Money	Commercial Paper Rate	AAA Rate	Price Expectations	$\bar{R}^2$	F	RMSE
	Coefficients							
3.1 1957.2-1973.4	.002 (3.1)	.781 (7.9)	-.626 (2.9)		-.625 (1.5)	.993	2183.6	5.57
3.2 1965.3-1973.4	.036 (2.1)	.589 (3.4)	-.260 (.7)		-1.235 (1.7)	.968	208.7	6.74
3.3 1957.2-1973.4	.029 (3.1)	.607 (4.9)		.037 (.1)	-.027 (.04)	.992	1942.3	7.38
3.4 1965.3-1973.4	.033 (1.3)	.334 (1.9)		.951 (.787)	-4.690 (2.03)	.996	197.9	4.18
3.5 1957.2-1973.4	.014 (4.3)	.904 (17.8)		-.611 (3.1)	-.825 (3.5)	.993	2559.7	3.44
3.6 1965.3-1973.4	.039 (2.2)	.546 (3.1)		-.219 (.641)	-.167 (1.9)	.968	208.8	6.75

Note: The constants are not reported. Values in parentheses are t-statistics. All models are estimated in level form. The process describing price expectations formation assumes T=20 quarters in 3.1 - 3.4 and T=10 in 3.5 and 3.6. The RMSE values are calculated for static forecasts of levels of money balances.

Table 4

Money Demand Estimates with Rational Price Expectations

Coefficients

Model	Constant	Real Income	Lagged Money	Commercial Paper Rate	Time Deposit Rate	π_1	π_2	$\bar{R}^2$	Standard Error	D.W.
4.1	.404	.135	.768	-.019	-.031					
	(2.56)	(4.49)	(12.46)	(7.29)	(2.79)			.987	.00472	1.87
4.2	2.903	.011	.965	-.952						
	(.69)	(6.67)	(42.31)	(7.11)				.987	1.0930	1.93
4.3	-14.17	.009	1.05	-.619		-.791		.986	1.0461	1.95
	(1.92)	(5.43)	(27.29)	(3.54)		(2.99)				
4.4	-4.48	.011	1.002	-.856			.295	.986	1.0916	1.92
	(.56)	(6.28)	(24.42)	(5.35)			(1.11)			
4.5	-6.47	.010	1.012	-.776		-.423		.988	.9521	1.92
	(1.43)	(6.43)	(41.74)	(5.94)		(5.28)				
4.6	7.76	.011	.939	-1.01			.161	.984	1.0818	1.91
	(1.52)	(6.41)	(34.47)	(6.91)			(1.78)			
4.7	-.511	.041	1.049	-.014		-.004		.988	.00463	1.95
	(3.13)	(6.41)	(30.86)	(4.66)		(3.63)				
4.8	-.005	.164	.583	-0.103		-.004		.387	.0053	2.21
	(.69)	(2.42)	(5.59)	(2.14)		(3.24)				
4.9	-.006	.153	.72	-0.13			-.004	.541	.0052	2.06
	(.87)	(2.46)	(7.65)	(3.05)			(3.29)			

Note: Values in parentheses are t-statistics. Models 4.1 and 4.7 are in log form, models 4.2 - 4.6 are in level form, and models 4.8 and 4.9 are first differences of log levels. Models 4.5 and 4.6 use the actual values of π; the others use <u>predicted</u> values of π from the first stage regressions.

Table 5

Summary of Static Post-Sample Forecast Error Tests

Model	Standard Error	RMSE (log-level)	RMSE (level)	Fraction of Error Due to Bias	No. Periods Overpredicted	Ratio of RMSE to Standard Error
5.1	.00472	.0207	4.7138	.8998	16	4.39
5.2	1.0930		2.6298	.7943	16	2.41
5.3	1.0461		1.2365	.0567	10	1.18
5.4	1.0916		2.0260	.6833	15	1.86
5.5	.9521		1.5257	.4723	13	1.60
5.6	1.0818		2.7412	.8506	16	2.54
5.7	.00463	.0048	1.0831	.0138	9	1.04
5.8	.0053	.0075	2.920	.2289	15	1.42
5.9	.00521	.0072	2.027	.1662	12	1.39
5.10	.0043	.0044	1.0011	.0104	9	1.02
5.11	.0043	.0049	1.1071	.0565	7	1.14

Note: The standard error is for the estimates reported in Table 4. The RMSE in column 2, rows 5.8 and 5.9, are for Δ ln forecasts. The forecasts are over 16 quarters beginning in 1974.1. Estimated coefficients for rows 5.10 and 5.11 are discussed in the text.