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MONETARY POLICY AND SHORT-TERM INTEREST RATES: AN EFFICIENT MARKETS-RATIONAL EXPECTATIONS APPROACH

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ABSTRACT

The impact of a money stock increase on nominal short-term interest rates has been a hotly debated issue in the monetary economics literature. The most commonly held view--also a feature of most structural macro models--has an increase in the money stock leading, at least in the short-run, to a decline in short interest rates. Monetarists dispute this view because they believe that it ignores the dynamic effects of a money stock increase.

This paper is an application of efficient markets-rational expectations theory to analyze empirically the relationship of money supply growth and short-term interest rates. This approach has the advantage over earlier research on this subject in that it imposes a theoretical structure that allows easier interpretation of the empirical results as well as more powerful statistical tests. In the interest of ascertaining the robustness of the results, many different empirical tests are carried out in this paper, and they uniformly do not support the proposition that increases in the money supply are correlated with declines in short rates.

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I

INTRODUCTION

The relationship between money supply growth and nominal interest rates is a hotly debated issue in the literature.[1] One view, associated with "Keynesian" structural macro models, has an increase in the money stock leading, at least in the short and medium runs, to a decline in interest rates.[2] An alternative view, associated with Milton Friedman (1968, 1969), indicates that interest rates might rise in response to an increase in money growth because the increase in money growth might lead to a rise in inflationary expectations and hence a rise in interest rates through a Fisher (1930) effect.

Previous empirical work on this issue has ignored constraints implied by the view that financial markets display rational expectations and are thus "efficient." Financial market efficiency should not be ignored because evidence supporting it is quite strong and recent work indicates that a failure to impose financial market efficiency on macroeconometric models can lead to highly misleading results.[3] In addition, a failure to impose the efficient markets (or, equivalently, rational expectations) constraints leads to a larger number of parameters to be estimated in this empirical work, and this leads to statistical tests with low power.[4]

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[1]Unless otherwise noted, whenever the phrase "interest rates" is used in this paper, it refers to nominal interest rates.

[2]For example, see Modigliani (1975).

[3]See Fama (1970) and Mishkin (1978).

[4]A more extensive discussion of the previous empirical work on this topic and references to this work can be found in Mishkin (1981).

The theory of efficient capital markets and rational expectations suggests an alternative approach for analyzing the relationship of money stock increases and interest rate movements. A previous paper (Mishkin (1981a) developed an efficient markets model of long interest rate determination, and then estimated this model using postwar quarterly data. This approach had the advantage of imposing a theoretical structure on the problem that allowed easier interpretation of the empirical results as well as more powerful statistical tests of the proposition that increases in the money stock are correlated with declines in long rates. In addition, a Keynesian, liquidity preference view of interest rate determination was embedded in the efficient markets model and tested.

This paper is a sequel to the earlier paper in that it conducts a similar analysis for short-term interest rates. The next section develops a rational expectations (or, equivalently, efficient markets) model for analyzing movements in short rates, and this model is estimated in the subsequent section. This paper then concludes with an interpretation of these results.

II

THE MODEL

The theory of rational expectations (or, equivalently, efficient markets theory) indicates that interest rates in a bond market should reflect all available information. To be more precise, it implies that the market uses available information correctly in assessing the probability distribution of all future interest rates and hence:

(1) $$E_m(r_t|\phi_{t-1}) = E(r_t|\phi_{t-1})$$

where

r_t = short-term (one period) interest rate at time t.

ϕ_{t-1} = information available at time t-1.

$E(\ldots|\phi_{t-1})$ = the expectation conditional on ϕ_{t-1}

$E_m(\ldots|\phi_{t-1})$ = the market's expectation (unbiased forecast) assessed at t-1.

If we denote the market's one-period-ahead forecast of the short rate of r_t^e (i.e., $r_t^e = E_m(r_t|\phi_{t-1})$) then (1) implies

(2) $$E(r_t - r_t^e|\phi_{t-1}) = 0$$

Equation (2) above states that the forecast error for short rates should be uncorrelated with any information or linear combinations of information in ϕ_{t-1}. An equivalent characterization of the rational expectations model which satisfies (2) is thus:

(3) $$r_t - r_t^e = (X_t - X_t^e)\beta + \varepsilon_t$$

where superscript e continues to denote the market's expectations conditional on all past available information and

X_t = a variable (or vector of variables) relevant to the determination of short-term interest rates,

β = a coefficient or vector of coefficients,

ε_t = serially uncorrelated error process (because $E(\varepsilon_t | \phi_{t-1}) = 0$).

The rational expectations model (3) stresses that an unanticipated change[5] in the short rate will occur only when unanticipated information hits the market. This distinction between the possible effects from unanticipated versus anticipated movements in variables is indeed an important feature of recent empirical work (for example, Barro (1977, 1978)).

In order to make the rational expectations model above empirically testable we must have a model of market equilibrium. Here we assume, as in Fama (1976b), that the one-period-ahead forward rate equals the one-period-ahead expected short rate plus a risk (liquidity) premium which varies with the uncertainty in short rate movements. I.e.:

(4) $F_t = r_t^e + \delta_t$

and

(5) $\delta_t = a_o + a_1 \sigma_t$

where

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[5] Since the anticipated change in the short rate equals $r_t^e - r_{t-1}$, $r_t - r_t^e$ is equivalent to the unanticipated change in the short rate.

F_t = forward rate for the one-period-rate at time t, implied by the yield curve at t-1,[6]

δ_t = risk (liquidity) premium for F_t.

σ_t = measure of uncertainty in short rate movements.

Combining the model of market equilibrium with (3), we have the rational expectations model estimated in this paper:

$$(6) \quad r_t - F_t = -a_o - a_1\sigma_t + (X_t - X_t^e)\beta + \varepsilon_t$$

As Fama (1976a) and Nelson and Schwert (1977) make clear, if the risk premium, δ_t, has small variation relative to other sources of variation in $r_t - F_t$, then the model of market equilibrium is not critical to empirical tests of the equation (6) model.[7] Although this type of situation frequently exists, making tests of financial market efficiency easy,[8] this is not the case here. Using a measure of uncertainty similar to Fama's (1976b), the amount of variation in r_t-F_t attributable to the variation of the liquidity premium is statistically significant at the 1% level in the 1959-76 sample period used here.[9] The appropriateness of this model of

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[6]In the case of 90 day treasury which are used here, the forward rate at the end of a quarter is calculated as

$$F_t = 4\{1-\frac{(360-180 \text{ rsix }_{t-1})}{(360-90\ r_{t-1})}\}$$

where all rates are in fractions and
$rsix_{t-1}$ = six month (180 day) bill rate at end of previous quarter,
r_{t-1} = 90 day bill rate at end of previous quarter.

[7]A more precise wording of this point would state that, in this case, tests of hypotheses concerning the model of the liquidity premium would have low statistical power.

[8]See for example Fama (1976a), Nelson and Schwert (1977) and Mishkin (1978, 1981a,b).

[9]See the results in equation (10) in the next section.

market equilibrium is thus an important factor that needs to be discussed further when the results of the (6) model are analyzed.

The research question posed in the Introduction suggests that the relationship of money growth and unanticipated changes in the short rate in a rational expectations model is of particular interest.[10] Substituting money growth for X_t in equation (6) leads to:

$$(7)\quad r_t - F_t = -a_o - a_1\sigma_t + \beta_m(MG_t - MG_t^e) + \varepsilon_t$$

where,

MG_t = the money growth rate at time t.

Thus, if unanticipated increases in the money stock are to have a negative correlation with unanticipated changes in short rates (as might be expected from "Keynesian" macro-econometric models), this implies that the coefficient on unanticipated money growth should be significantly negative in equation (7): i.e., $\beta_m < 0$.

An important caveat is in order. The rational expectations model does not guarantee that equation (3) is a reduced form where $X_t - X_t^e$ is exogenous so that the estimates of β are consistent.[11] It implies only that $r_t - r_t^e$ is correlated with unanticipated movements in variables. Another way of stating this point is to acknowledge that the rational expectations model does not indicate whether a significant β coefficient implies causation from its unanticipated variable to short-term interest rates. Regarding rational expectations, causation could run in the other direction, or it could be nonexistent as in the case where new information is simultaneously affecting both unanticipated

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[10] As has been found in foreign exchange markets (See for example Mussa (1979)) quarterly changes in the spot rate, in this case of the short rate, are primarily attributable to unanticipated movements in the spot rate. See Fama (1976b). Using the model of the liquidity premium estimated in (10), the correlation of unanticipated short rate movements, and the actual change in short rates is high in the 1959-76 sample period used here, being greater than .8. Thus, results obtained in this study for unanticipated changes in short rates also apply to changes in short rates.

[11] This issue of the consistency of the β estimates is discussed more [illegible]

short rates and the right-hand-side variable. Thus, we must be careful in interpreting empirical results on the β's not to ascribe causation to the results without further identifying information.

The above caveat must be kept in mind especially when we analyze the estimated β_m coefficient. If the money supply process is seen as exogenous -- a view that has received some support in the literature[12]--the interpretation of the estimated β_m is straightforward. The finding of a significant negative β_m would then provide evidence supporting the "Keynesian" position that increased money growth will, at least in the short-run, lead to declines in short rates; and a failure to find this result would cast doubt on this view. However, if the money supply process is not exogenous, the position taken by many critics of monetarist analysis, then the estimated β_m coefficient might suffer from simultaneous equation bias and give a misleading impression as to the effect of an increase in the money supply on short-term interest rates. Because the analysis in this paper provides no information on the exogeneity of the money supply process, the β_m estimates in the discussion of the empirical results are viewed only as providing information on the correlations of unanticipated money growth and the unanticipated change in short rates. Interpretation of these correlations is then deferred to the concluding remarks toward the end of the paper.

The liquidity preference approach to the demand for money suggests other relevant information which might be concluded in the X-vector of the

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[12]Sims (1972) contains a discussion of the differing views in the literature on the exogeneity of the money supply process and finds evidence which he interprets as supporting the view that causality runs from money growth to income rather than the other way around. As Jacobs, Leamer and Ward's (1979) and Zellner's (1979) criticisms of these causality tests indicate, however, these tests do not resolve the issue of the exogeneity of money supply process.

rational expectations model. Changes in interest rates are related not only to changes in the money stock but also to changes in real income, the price level and inflation.[13] Hence short-term interest rates might be related not only to the growth rate in the nominal money stock, as in equation (7), but also to the growth rate of real income and inflation. Adding this information to the X-vector in the equation (6) model leads to the following:

$$(8) \quad r_t - F_t = \alpha_o - \alpha_1\sigma_t + \beta_m(MG_t - MG_t^e) + \beta_y(YG_t - YG_t^e) + \beta_\pi(\pi_t - \pi_t^e) + \varepsilon_t$$

where

YG_t = growth rate of real income,

π = inflation rate,

β_m, β_y, β_π = coefficients.

This equation is really a rational expectations analog to the typical money demand relationship found in the literature. In addition, equation (8) captures elements of interest rate models of the Feldstein and Eckstein (1970) variety.

The money demand view of equation (8) indicates that the income coefficient, β_y, should be positive: i.e. $\beta_y > 0$. However, the signs of the unanticipated money growth and inflation coefficients are not as straightforward because they depend on the time-series process of money growth and inflation. A positive effect of an unanticipated increase in inflation ($\beta_\pi > 0$) follows from the resulting reduction in real money balances. The positive unanticipated inflation effect is further strengthened if the time-series process of inflation is such that, as in the Cagan (1956) adaptive expectations model, an unanticipated rise in inflation leads to a higher expected inflation rate in the coming period. Then a Fisher

[13] A more explicit demonstration of this point can be found in Mishkin (1981).

effect will lead to higher short rates. The more persistent the time-series process of inflation - - that is, the more an unanticipated increase in inflation leads to a continuing increase next period--the larger the Fisher effect and β_π should be. Since a surprise in money growth will affect short rates in part through the "price anticipations" effect, the β_m coefficient will also not be independent of the time-series process of money growth.

We now turn to the actual estimation of the rational expectations models of (7) and (8).

III

EMPIRICAL RESULTS

THE DATA

Six month treasury bills were not issued before 1959, and the six month bill rate is needed to calculate the forward rate used here. Thus, the empirical results below use postwar quarterly data over the 1959-76 sample period. The data sources and definitions of the variables used in these estimates are as follows:

r_t = the 90 day treasury bill rate, the last trading day in the quarter - - in fractions.

$M1G_t$ = growth rate of M1 (quarterly rate) = the first differenced series of the log of the average level of M1 in the last month of the quarter.

$M2G_t$ = growth rate of M2 (quarterly rate) = the first differenced series of the log of the average level of M2 in the last month of the quarter.

IPG_t = growth rate of industrial production (quarterly rate) = the first differenced series of the log of Industrial Production in the last month of the quarter.

π_t = the CPI inflation rate (quarterly rate) - the first differenced series of the log of CPI in the last month of the quarter.

UN_t = unemployment rate in the last month of the quarter - - in percent.

BOP_t = balance of payments on current account for that quarter - - in billions.

Unless otherwise noted, all these variables have been constructed from seasonally adjusted data except for r_t and F_t which do not require seasonal adjustment. The BOP variable was obtained from the NBER data bank, while the IPG, π , and UN variables were constructed from data in the Commerce Department's <u>Business Statistics</u> and <u>Survey of Current Business</u>. The M1 and

M2 data were obtained from the Board of Governors of the Federal Reserve *Banking and Monetary Statistics* and the *Federal Reserve Bulletin*, while the data for r_t and F_t were supplied by the Board of Governors of the Federal Reserve.

Since misleading results can be obtained from efficient markets models using averaged data,[14] the data for bond returns are derived from security prices at particular points in time. In keeping with this, an attempt has been made to derive the other variables used here with data as close to being end of quarter as possible. For this reason, Industrial Production is used as a proxy for real income in estimating equation (8) rather than a more broadly based National Income Accounts measure. Similarly, the CPI has been used to calculate the inflation variable rather than the GNP deflator.

THE ESTIMATION METHOD

In order to estimate the efficient markets models of equations (7) and (8), measures of anticipated money growth, income growth and inflation must be developed. Here, anticipations of variables in the information set X are assumed to be optimal linear forecasts using time-series models of the following form:

(9) $X_{i,t} = Z_{i,t-1}\,\gamma_i + u_t$

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[14]See Working (1960) for example.

[15]In previous work (Mishkin (1981), experiments with quarterly averaged data led to substantially worse fits for equations similar to (7) and (8), fewer significant coefficients and no appreciable differences as to the statistical significance of the β_m coefficients.

where

X_i = MG, IPG, or π,

$Z_{i,\ t-1}$ = a vector of variables containing information available at time t-1 - - this includes variables known before t-1 as well as at t-1,

u_t = white noise error term ,

γ_i = a vector of coefficients,

and the subscript i refers to either MG, IPG, or π.

A critical issue in the research strategy used here is the methodology for choosing the specification of the time-series models of (9). It is difficult theoretically to exclude any particular piece of information available at time t-1 as a useful predictor of an $X_{i,t}$ variable. For example, economic theory cannot provide much guidance as to which variables to exclude in a money growth equation. Even though there is no strong theoretical reason for expecting a particular variable to enter the Z-vector, it might be a useful predictor of money growth because the personalities involved in policymaking could be such that they react to this variable for their own inscrutable reasons. Thus the theoretical model a researcher uses to explain this money growth specification might be relatively unimportant in deciding the validity of his particular specification versus that of another researcher.

The discussion above suggests that an atheoretical statistical procedure might be superior to economic theory for deciding on the specification of the time-series models in (9). Furthermore, because theory is less of a useful guide in evaluating the time-series models needed here than is true in other empirical work, it is more important to check for the robustness of results by using several model specifications in estimating the rational expectations model. In keeping with this line of thinking, two procedures for specifying the time-series models of (12) are used in the text, with several additional specifications in the results discussed in the Appendix.

The simplest equations which can be used to describe money growth, industrial production growth and inflation are univariate time-series models of the autoregressive type. Fourth order autoregressions are usually successful in reducing quarterly data's residuals to white noise and are thus used here. The resulting estimates for M1G, M2G, IPG and π can be found in Table 1. Note that there is a fair amount of persistence in the time-series models for money growth and inflation, indicating that "price anticipation" effects of the sort that Friedman (1968, 1969) discusses are potentially important.

More complex multivariate time series models have been estimated using the following procedure. Each of the four variables - - M1G, M2G, IPG, and π - - was regressed on its own four lagged values as well as on four lagged values of each of the other three variables and four lagged values of each of the following variables: the unemployment rate; the 90 day treasury bill rate; the balance of payments on current account; the growth rate of real federal government expenditure, the high employment budget surplus, and the growth rate of federal government, interest bearing debt, in the hands of the public.[16] (These other variables were selected because a reading of the literature on Federal Reserve reaction functions indicated that they might help explain money growth.[17]) The four lagged values of each variable were

[16]The source of these variables in the NBER data bank.

[17]See Fair (1978) and the references therein.

TABLE 1

UNIVARIATE TIME-SERIES MODELS

Model No.	1.1		1.2		1.3		1.4	
	Dependent Variable							
Coefficient of	M1G		M2G		IPG		π	
Constant term	.0053 (.0019)		.0076 (.0024)		.0097 (.0035)		.0014 (.0008)	
M1G(-1)	.3158 (.1234)	F(4,67) = 4.25						
M1G(-2)	.2166 (.1289)							
M1G(-3)	.0306 (.1275)							
M1G(-4)	-.0371 (.1236)							
M2G(-1)			.6113 (.1220)	F(4,67) = 10.13				
M2G(-2)			-.0412 (.1421)					
M2G(-3)			.1619 (.1408)					
M2G(-4)			-.1525 (.1230)					
IPG(-1)					.3514 (.1187)	F(4,67) = 2.94		
IPG(-2)					-.2100 (.1250)			
IPG(-3)					.1449 (.1238)			
IPG(-4)					-.2025 (.1131)			
$\pi(-1)$							.3991 (.1209)	F(4,67) = 51.77
$\pi(-2)$							.6162 (.1268)	
$\pi(-3)$							-.0179 (.1275)	
$\pi(-4)$							-.1613 (.1112)	
R^2	.2023		.3766		.1496		.7555	
Standard Error	.0066		.0070		.0239		.0038	
Durbin-Watson	1.96		1.94		1.98		2.01	

Note: Standard errors of the coefficients are in parentheses and the F-statistics test the joint null hypothesis that the coefficients are equal to zero. Note that because lagged dependent variables appear in the time-series models above, all the test statistics are only valid asymptotically.

retained in the equation only if they were jointly significant at the five percent level or higher. The major advantage of this procedure is that it imposes a discipline on the researcher that prevents his searching for model specifications that confirm his priors.

The resulting multivariate time-series models can be found in Table 2, along with F-statistics of the joint significance test for whether the four lagged values of each variable should be included in the regression model.[18, 19] Note that these multivariate time-series models contain some information of independent interest because they make use of Granger's (1969)

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[18]Note that the because lagged dependent variables appear in the time-series models of Tables 1 and 2, all the test statistics are only valid asymptotically. The F-statistics which test the null hypothesis that the lagged values of the following variables add no explanatory power to the Table Table 2 regressions are as follows. In the 2.1 regression, the $F(4, 59)$ statistics are: 1.81 for π, 1.26 for IPG; 1.42 for UN; .47 for the growth rate of real government expenditure; .79 for the balance of payments; .62 for the high employment surplus; .43 for the growth rate of government debt, and 2.46 for r. In the 2.2 regression, the $F(4, 59)$ statistics are .40 for π; 2.17 for M1G; 1.65 for IPG; .25 for UN; .22 for the growth rate of real government expenditure; .70 for the balance of payments on current account; 1.28 for the high employment surplus; and .44 for the growth rate of government debt. In the 2.3 regression, the $F(4, 47)$ statistics are: .46 for the growth rate in real government expenditure; .09 for M1G; .70 for the high employment surplus; 1.84 for the growth rate of government debt; and 2.32 for UN. In the 2.4 regression, the $F(4, 51)$ statistics are: 1.13 for M2G; .94 for r; 1.55 for IPG; 1.59 for the growth rate in real government expenditure; 2.01 for the high employment surplus; and .33 for the growth rate of government debt.

[19]Chow (1960) tests where the sample has been split into equal halves reveal that there is some instability in the coefficients for both the univariate and multivariate M1 money growth model. The Chow test for the model of 1.1 yields $F(5, 62) = 2.73$, while for 2.1 $F(9, 54) = 2.94$, both of which are significant at the 5% level. However, neither of the M2 money growth models displays this instability. For model 1.2 $F(5, 62) = 1.24$ and for 2.2 $F(9, 54) = 1.80$. Of the IPG and π models only the univariate π model displays coefficient instability: For model 1.3 $F(5, 62) = 3.80$, for 1.2 $F(5, 62) = 1.24$, for 2.2 $F(9, 54) = 1.80$, for 1.4 $F(5, 62) = 1.83$ and for 2.4 $F(17, 38) = 1.24$.

TABLE 2

MULTIVARIATE TIME-SERIES MODELS

Model No.	2.1		2.2		2.3		2.4
	Dependent Variable						
Coefficient of	M1G		M2G		IPG		π
Constant term	.0015		.0044		.0017		.0033
	(.0021)		(.0027)		(.0111)		(.0021)
M1G(-1)	-.1031						.1835
	(.1875)						(.0635)
M1G(-2)	.5336						.0376
	(.1906)	$F(4,63)$					(.0649)
M1G(-3)	-.3184	= 2.44					-.0663
	(.1990)						(.0657)
M1G(-4)	.0052						.1600
	(.1739)						(.0727)
M2G(-1)	.5998		.5211		.7132		
	(.1666)		(.1271)		(.3814)		
M2G(-2)	.5612		.0955		.9460		
	(.1935)	$F(4,63)$	(.1426)	$F(4,63)$	(.4300)	$F(4,51)$	
M2G(-3)	.4112	= 5.28	.1179	=10.42	.7217	=11.09	
	(.2077)		(.1325)		(.4122)		
M2G(-4)	.0240		-.0838		-.2049		
	(.1843)		(.1180)		(.4182)		
IPG(-1)					-.2778		
					(.1129)		
IPG(-2)					-.3626		
					(.1089)	$F(4,51)$	
IPG(-3)					.0457	= 4.70	
					(.1089)		
IPG(-4)					-.1869		
					(.1010)		
$\pi(-1)$					-2.277		.1209
					(.793)		(.1273)
$\pi(-2)$					-1.753		.7705
					(.701)	$F(4,51)$	(.1171)
$\pi(-3)$					.398	= 5.05	.3911
					(.846)		(.1339)
$\pi(-4)$					1.252		-.2328
					(.591)		(.1176)
r(-1)			-.5214		.9813		
			(.1221)		(.4501)		
r(-2)			.4827		.5253		
			(.1631)	$F(4,63)$	(.4819)	$F(4,51)$	
r(-3)			.0551	= 5.93	.3307	= 5.62	
			(.1669)		(.4793)		
r(-4)			.0333		-1.830		
			(.1279)		(.4368)		

TABLE 2 (cont.)

Model No.	2.1	2.2	2.3		2.4
	Dependent Variable				
Coefficient of	M1G	M2G	IPG		π
UN(-1)					-.0052
					(.0014)
UN(-2)					-.0016
					(.0021)
UN(-3)					.0034
					(.0021)
UN(-4)					-.0004
					(.0013)
BOP(-1)			.0058		.0001
			(.0035)		(.0006)
BOP(-2)			.0050		.0003
			(.0047)	$F(4,51) = 3.54$	(.0008)
BOP(-3)			.0116		-.0005
			(.0046)		(.0008)
BOP(-4)			-.0094		-.0010
			(.0037)		(.0006)
R^2	.4024	.5472	.7108		.8789
Standard Error	.0059	.0061	.0160		.0029
Durbin-Watson	2.01	1.95	2.00		2.21

Note: Standard errors of the coefficients are in parentheses and the F-statistics test the joint null hypothesis that the coefficients are equal to zero. Note that because lagged dependent variables appear in the time-series models above, all the test statistics are only valid asymptotically.

concept of predictive content.[20] One interesting feature of the multivariate money growth equations is that, in contrast with Barro's (1977) work, no fiscal policy or unemployment variables were found to be statistically significant at the five percent level.[21]

Before turning to the procedures for estimating the rational expectations model, the measure of short rate uncertainty (σ_t) used here requires some discussion. Fama (1976b) calculates σ_t as the average of the absolute values of the changes in the spot rate during the year before t and during the year following t. Because the risk (liquidity) premium must be set conditional on available information - - in this case that known at t-1 - - allowing σ_t to be calculated from information not available at t-1 does pose some conceptual difficulties. An alternative, though similar, measure of σ_t is used in this study. The difference between the forward rate and the spot rate, i.e., $r_t - F_t$, was regressed on measures of σ_t, calculated as the average absolute change of the bill rate over a number of previous quarters, where the number of quarters was varied. The best fit was obtained with σ_t calculated from eight previous quarters of changes in the bill rate. The results are as follows

$$(10)\quad r_t - F_t = \underset{(.0017)}{-.0001} - \underset{(.2937)}{1.0961}\,\sigma_t + \varepsilon_t$$

$R^2 = .1659$ Standard Error $= .0068$ Durbin-Watson $= 1.90$

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[20] I prefer not to refer to this as Granger-causality here because this nomenclature has led to much confusion in the literature. For a discussion of this point see Zellner (1979).

[21] Sheffrin (1979) also finds that fiscal policy variables do not help explain money growth.

where

$$\sigma_t = \frac{\sum_{L=1}^{8} | r_{t-i} - r_{t-i-1} |}{8} .$$

As in Fama (1976b), increased uncertainty in short rate movements does lead to an increased risk premium and this effect is statistically significant at the 1% level. In addition, the σ_t measure used here outperforms the Fama measure that is constructed from information unavailable at t-1. The above measure of σ_t is used in the empirical tests that follow. However, its specification is not a critical issue to the outcomes: use of a Fama measure of σ_t or the exclusion of σ_t from the model altogether does not alter the results appreciably.

One way to proceed in estimating the rational expectations model is to use a two-step procedure outlined in the recent work of Barro (1977, 1978). After estimating the time-series models of Tables 1 and 2, the residuals from these regressions can be used as proxies for the corresponding unanticipated variables in estimating equations (8) and (11). Tests of whether only unanticipated changes are related to $r_t - r_t^e$ could then also proceed as in Barro (1977).

Although the empirical results using the above approach are not unreasonable and are similar to those produced here in the text (see the Appendix), there are serious econometric criticisms of this approach.[22] However, an econometric technique which does not suffer from these criticisms is outlined below.[23]

[22] See Mishkin (1980).

[23] The technique used here is quite similar in concept to that proposed by Sargent (1979), although it is somewhat easier to execute and notationally simpler.

In matrix notation, the rational expectations model can be written as:

$$(11)\quad r - F = -a_o - \sigma\, a_1 + (X - X^e)\,\beta + \varepsilon$$

where

$r - F = n \times 1$

$a_o = n \times 1$

$\sigma = n \times 1$

$a_1 = 1 \times 1$

X and $X_e = n \times k$

$\beta = k \times 1$

$\varepsilon = n \times 1$

n = number of observations

k = number of right-hand-side variables in X.

The linear time series model for X, whether univariate or multivariate can also be written as:

$$(12)\quad X = Z\gamma + U$$

where

Z = an $n \times m$ matrix of lagged variables, where m is the number of variables,

γ = $m \times k$ matrix of coefficients,

U = $n \times k$ matrix of white noise error terms.

The optimal linear forecast of X is then

$$(13)\quad X^e = Z\gamma$$

and substituting this into (11) we have:

$$(14)\ r - F = -a_o - \sigma a_1 + (X - Z\gamma)\,\beta + \varepsilon$$

The system in (12) and (14) can be stacked into one regression system with n(k+1) observations, and it can be estimated by non-linear least squares methods imposing the constraints that the γ in (12) and (14) are equal. In order to obtain more efficient parameter estimates as well as consistent test statistics, corrections must be made for heteroscedasticity both within and across equations in this system.[25]

This procedure is superior to the alternative two-step procedure.[26] More efficient parameter estimates of β and γ will result because both (12) and (14) make use of information from each other in the estimation process. In addition, it generates a simple test of the model which is similar to recent tests of "rationality" in the literature that proceed along the lines of Modigliani and Shiller (1973).[27] It is a simple like-

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[25]The following iterative procedure was used to correct for heteroscedasticity in these estimates. In the first stage estimation of the non-linear system, if Goldfeld-Quandt tests indicated that heteroscedasticity existed within an equation, the variables in this equation were weighted using a time trend procedure outlined by Glesjer (1969). Furthermore, the variables in each equation of the system were appropriately weighted so that each equation individually had the same sum of squared residuals. After the first stage estimation, the sum of squared residuals for each equation were calculated and were then used to weight the variables in each equation so that the sum of squared residuals were the same in all cases. Then the non-linear system was estimated all over again. This resulted in a similar sum of squared resiudals in all equations of the (12) and (14) system so that no further iterations were performed. Some experimentation did indicate that no appreciable differences in the results occured if some modification of this iterative procedure was used which left the sum of squared residuals reasonably equal for all the equations. For further details on this estimation procedure, see Mishkin (1980).

[26]See Abel and Mishkin (1980) for a more detailed discussion of the econometric issues in using the joint non-linear procedure.

[27]See Pesando (1975), Sargent (1979), Carlson (1977), Mullineaux (1978), Freidman (1978), and Mishkin (1981b).

lihood ratio test for whether the (12) and (14) system satisfies the non-linear constraints implied by the equality of γ in (12) and (14). The likelihood ratio statistic, $-2 \log(L^c/L^u)$, is distributed asymptotically as $\chi^2(q)$ where q is the number of non-linear constraints

L^c = likelihood of the estimated constrained system,

L^u = likelihood of the estimated unconstrained system.

In this non-linear least squares system, the likelihood ratio statistic is

$$n(k + 1)(\log(SSR^c) - \log (SSR^u))$$

where

SSR^c = sum of squared residuals from the constrained system,

SSR^u = sum of squared residuals from the unconstrained system.

THE RESULTS

Because there is no strong theoretical reason for estimating the rational expectations model with one monetary aggregate versus another, unanticipated growth rates of both M1 and M2 are used in estimation. The resulting estimates and test statistics of this data appear in Table 3. Panel A of this table contains estimates where only univariate models of the form found in Table 1 are used in (12), while Panel B uses the multivariate models of the form found in Table 2. The estimates of the γ coefficients are not presented here because they are similar to those found in Tables 1 and 2.

The first issue we should look at is whether the non-linear constraints implied by the model are satisfied. The likelihood ratio tests reported in Table 4 indicate that they are not. The marginal significance levels in Table 4 are the probability of obtaining that value of the likelihood ratio statistic or higher under the null hypothesis that the non-linear constraints are valid. They indicate that the constraints are rejected at the 5% level in six out of eight cases. How should we interpret these rejections?

TABLE 3

NON-LINEAR ESTIMATES OF THE EFFICIENT-MARKETS--RATIONAL EXPECTATIONS MODEL USING SEASONALLY ADJUSTED DATA

Model No.	Coefficients of					
	$(M1G-M1G^e)$	$(M2G-M2G^e)$	$(IPG-IPG^e)$	$(\pi-\pi^e)$	constant term	σ
			Panel A. Using Univariate Models in (15)			
3.1	.2788* (.1088)				.0006 (.0015)	-1.2266** (.2714)
3.2	.2774** (.1075)		.0352 (.0275)	.6211** (.1989)	.0002 (.0014)	-1.1514** (.2618)
3.3		.1616 (.1085)			.0006 (.0015)	-1.2563** (.2851)
3.4		.1904 (.1053)	.0399 (.0278)	.6545** (.2058)	.0002 (.0015)	-1.1571** (.2686)
			Panel B. Using Multivariates Models in (15)			
3.5	.1677 (.1283)				.0006 (.0015)	-1.2761** (.2863)
3.6	.2512 (.1381)		-.0455 (.0493)	.5199 (.3272)	.0004 (.0016)	-1.2109** (.3015)
3.7		.2562 (.1341)			.0001 (.0016)	-1.1807** (.2917)
3.8		.3039* (.1409)	-.0770 (.0471)	.6501* (.3314)	-.0004 (.0016)	-1.0779** (.3069)

Note: * = significantly different from zero at the 5 percent level.
** = significantly different from zero at the 1 percent level.
Asymptotic standard errors of the coefficients are in parentheses.

TABLE 4

LIKELIHOOD RATIO TESTS OF NON-LINEAR CONSTRAINTS

Model No.	Likelihood Ratio Statistic	Marginal Significance Level
3.1	$\chi^2(4) = 12.76$	.0125
3.2	$\chi^2(12) = 13.65$	.3235
3.3	$\chi^2(4) = 12.46$	.0143
3.4	$\chi^2(12) = 17.18$	.1430
3.5	$\chi^2(8) = 21.65$	.0056
3.6	$\chi^2(28) = 50.02$	.0064
3.7	$\chi^2(8) = 25.69$	.0012
3.8	$\chi^2(28) = 50.92$	.0051

Note: The marginal significance level is the probability of getting that value of the likelihood ratio statistic or higher under the null hypothesis.

The non-linear constraints are generated by two hypotheses: 1) rational expectations and 2) the model of market equilibrium in equation (6). A rejection of these constraints could thus result from the failure of either hypothesis. In the situation where use of the appropriate model of market equilibrium is unimportant in the test results because it contributes so little variation to the variable of interest, then rejections of the non-linear constraints indicate that expectations are not rational. In this case the rationale for the analysis of this paper would disappear. However, as discussed in the previous section, the contribution of the model of market equilibrium to the variation of $r_t - F_t$ appears to be large. A rejection of the non-linear constraints is then likely to result from a poor specification of this market equilibrium model.

There is a substantial body of evidence supporting the rationality of expectations in bond markets,[28] and this leads to a suspicion that it is the model of market equilibrium that causes the rejections in Table 4.[29] Fortunately, if this is the source of the rejection, the rational expectations model estimated here is still a valid framework for analyzing the relationship of money growth and short interest rates. With rational expectations, the unanticipated $X_t - X_t^e$ variables will be uncorrelated with any past information, among which can be included the determinants of the risk premium which is set at t-1. Therefore, if some determinants of this risk premium have been excluded from the market equilibrium model, with the resulting rejection of the non-linear constraints, this will not lead to

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[28] See for example, the survey in Fama (1970) as well as more recent work such as Mishkin (1978, 1981a, b) and Sargent (1979).

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[29] With a suitable transformation of the unconstrained system outlined in Abel and Mishkin (1980), additional evidence is available on the potential misspecification of the model of market equilibrium. The unconstrained system where the γ are not equal in (12) and (14) can be rewritten as

$$X = Z\gamma + u \tag{12'}$$

$$r-F = -a_0 - \sigma a_1 - Z\alpha + (X - Z\gamma)\beta + \varepsilon \tag{14'}$$

where the γ's are constrained to be equal in (12') and (14'). Therefore, the non-linear constraints tested in this paper are equivalent to $\alpha = 0$ in the above system. It is now easy to see the following point: If the liquidity premium is related to the variables in Z yet they have been excluded from the model of the liquidity premium, then this could explain the rejections of the non-linear constraints found here. To make this conjecture plausible, we should expect that a model of the liquidity premium which is related to Z would have reasonable characteristics. For example the Fama-type model of the liquidity premium in equation (10) does generate plausible values. The resulting liquidity premiums (at annual rates) have a mean of 57 basis points and a standard deviation of 30 basis points. They also move smoothly: their autocorrelations for lags of one through four quarters are respectively .96, .91, .85, and .78. In the model which leads to the strongest rejection of the non-linear constraints, model 3.7, we could attribute this rejection to the fact that a more appropriate specification of the liquidity premium is $\delta = a_0 + a_1\sigma + Z\alpha$, where Z contains the four lagged values of money growth (M2G) and treasury bill rates (r). This latter specification leads to values for the liquidity premium that are somewhat more variable and less smooth than the equation (10) specification, but not appreciably so. The liquidity premiums from this expanded specification have a mean of 57 basis points, a standard deviation of 46 basis points and four lagged autocorrelations of .75, .56, .49 and .29.

Viewing the rejections with the benefit of the system (12') and (14') also has the advantage that it provides us with potentially interesting information on the liquidity premium. The results in Tables 4 and 6 indicate that the liquidity premium could be related to money growth and interest rates as well as the variability measure σ. However, they give no indication that the liquidity premium is in addition related to the other variables in Table 2; π, IPG, UN and BOP. The results here thus point out a direction for future research on the liquidity premium. Following Nelson (1972), I also conducted more direct experiments on the relation of the liquidity premium to lagged r and UN with negative results. Experiments with lagged values of r-F also did not add explanatory power to the model of the liquidity premium.

inconsistent estimates of the β coefficients.[30] Since the derivation of a better model of the risk premium is not necessary for acheiving reliable estimates of the β's, this tricky research issue, which is beyond the scope of this paper, is left as a subject for future research.

The unanticipated M1G coefficients in Panel A do not support the view that an unanticipated increase in money growth is correlated with an unanticipated fall in short rates. Not only are both of these coefficients in model 3.1 and 3.2 positive rather than negative, but they are also significantly different from zero. The coefficients are not numerically small either. They indicate that a 1% surprise increase in M1 is associated with a 28 basis point unanticipated increase in the bill rate.[31] The Panel B estimates of the M1G coefficients indicate that the above conclusion on the relationship of short rates and M1 growth is not altered as a result of using multivariate versus univariate time-series models to describe expectations formation. Again both coefficients are positive, although in this case neither is significantly different from zero.

How different are the results found here from those that might be inferred from "Keynesian" structural macro-econometric models? The response of one such model, the MPS (MIT-PENN-SSRC) Quarterly Econometric Model (1977), to a 1% surprise increase in M1 growth was analyzed with a simulation technique discussed in Mishkin (1979)[32]. The MPS model indicated that this 1%

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[30]This depends on a proper specification of equation (15) so that $X - Z\gamma$ is uncorrelated with any past information used to set the risk premium.

[31]A basis point is defined as 1/100 of a percentage point. I.e., a one basis point rise in a 5% short rate would denote an increase to 5.01%.

[32]More details on this simulation experiment are given in Mishkin (1981a).

Ml surprise led to an immediate decline of 88 basis points in the bill rate. This strongly constrasts with the finding here that even the least positive Ml coefficient is more than five standard deviations away from this figure.

The similarity between the money growth as well as other coefficients estimates in going from Panel A to Panel B is encouraging for it gives us confidence that the results found here are robust to changes in the models describing expectations.[33] Additional results described in the Appendix also support this view. Note that the asymptotic t-statistics for the money growth and inflation coefficients in Panel A tend to be higher than those in Panel B, thus yielding stronger results. This lends some support to the position taken by Feige and Pearce (1976) that forecasts from univariate time-series models may be "economically rational" expectations.

The coefficients on unanticipated M2 growth tell a similar story to the the Ml growth coefficients. They also do not support the view that unanticipated money growth is associated with an unanticipated decline in short rates. The Panel A results for M2 are not as strong as the Ml results in supporting a positive correlation between unanticipated money growth and short rates: both M2 coefficients are positive, but neither is significantly different from zero. However, in Panel B, one of the positive M2G coefficients is statistically significant while this is not the case for the MlG coefficients.

The results on the unanticipated inflation and industrial production coefficients in Panel A do conform to our priors. In both the Ml and M2

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[33]As Feige and Pearce (1976) have argued, once past information on the variable to be forecast is used in forecasting, other information might have little incremental predictive power. The similarity of results in Panel A and B gives some credibility to this viewpoint and this issue is discussed more extensively in Mishkin (1980).

rational expectations models, these coefficients are positive and the inflation coefficients are significantly different from zero at the 1% level. The Panel B results for inflation are similar to those in Panel A, although now only the inflation coefficient in the M2 model is significant. The results on unanticipated industrial production growth continue to be weak in Panel B, and here these coefficient now have the "wrong" sign although they are insignificant.

The rational expectations model does not specify whether the $X - X^e$ variables should be described by seasonally adjusted rather than seasonally unadjusted data. This is an empirical issue that cannot be settled easily on theoretical grounds because it is not clear whether market participants concentrate on seasonally adjusted versus unadjusted information. For this reason, the (15) and (17) system has also been estimated with seasonally unadjusted data for the X's over the 1959-76 sample period. The resulting estimates and test statistics appear in Tables 5 and 6 and were obtained with the same techniques as the previous estimates with seasonally adjusted data.

A comparison of the Tables 3 and 4 with the Tables 5 and 6 results indicates that the use of adjusted versus unadjusted data is not a critical factor in this research. The likelihood ratio tests of the non-linear constraints in Table 6 have similar marginal significance levels to those in Table 4, and now five out of eight tests reject these constraints at the 5% level. In addition the coefficient estimates are similar to those in Table 4,and the Panel A money growth, inflation and industrial production growth coefficients have larger asymptotic t-statistics than those in Panel B.

There are two important differences in the adjusted versus the unadjusted results. The unadjusted industrial production coefficients now all

TABLE 5

NON-LINEAR ESTIMATES OF THE EFFICIENT-MARKETS--RATIONAL-EXPECTATIONS MODEL WITH SEASONALLY UNADJUSTED DATA

Model No.	Coefficients of					
	$(M1G - M1G^e)$	$(M2G-M2G^e)$	$(IPG - IPG^e)$	$(\pi-\pi^e)$	constant term	σ
		Panel A. Using Univariate Models in (15)				
5.1	.3029** (.0652)				.0003 (.0014)	-1.1255** (.2530)
5.2	.2458** (.0671)		.0274 (.0171)	.4687** (.1716)	.0001 (.0014)	-1.1267** (.2464)
5.3		.1926* (.0644)			.0003 (.0015)	-1.1468** (.2765)
5.4		.1967** (.0624)	.0440* (.0176)	.5459** (.1746)	.0001 (.0014)	-1.1260** (.2526)
		Panel B. Using Multivariate Models in (15)				
5.5	.3431** (.0831)				.0007 (.0015)	-1.2403** (.2639)
5.6	.2484** (.0956)		.0386 (.0376)	.5079 (.2623)	.0004 (.0016)	-1.2037** (.2861)
5.7		.3285** (.0918)			-.0007 (.0015)	-.9891** (.2841)
5.8		.2011* (.0986)	.0400 (.0374)	.5788* (.2674)	-.0003 (.0016)	-1.0791** (.3021)

Note: * = significantly different from zero at the 5% level.
** = significantly different from zero at the 1% level.
Asymptotic standard errors of the coefficients are in parenthesis.

Table 6

LIKELIHOOD RATIO TESTS OF NON-LINEAR CONSTRAINTS

Model #	Likelihood Ratio Statistic	Marginal Significance Level
5.1	$\chi^2(4) = 9.14$	.0578
5.2	$\chi^2(12) = 14.81$	.2521
5.3	$\chi^2(4) = 12.02$	.0172
5.4	$\chi^2(12) = 15.01$	.2407
5.5	$\chi^2(8) = 19.30$	.0133
5.6	$\chi^2(28) = 49.71$	.0070
5.7	$\chi^2(8) = 24.90$	.0016
5.8	$\chi^2(28) = 49.96$	.0065

Note: The marginal significance level is the probability of getting that value of the likelihood ratio statistic or higher under the null hypothesis.

have the expected positive sign in Table 5, and one of these coefficients is even statistically significant at the 5% level. Of even greater interest are the stronger results on the relationship of money growth and short rates when seasonally unadjusted data is used. Not only are all the coefficients on unanticipated money growth positive in Table 5, but seven of them are statistically significant at the one percent level and the remaining coefficient is significant at the 5% level.

The unadjusted data then provide much stronger evidence than the adjusted data that an unanticipated increase in money growth is not associated with an unanticipated decline in short rates, as we might expect from "Keynesian" macro-econometric models. Rather, the reverse seems to be the case.

IV

CONCLUDING REMARKS

A wide range of empirical tests exploring the relationship of money growth and short-term interest rates have been conducted in this paper and in the Appendix. A guiding principle in this research has been the use of many different empirical tests of the model in order to provide information on the robustness of the results. The pursuit of this goal has led to model estimation where there have been variations along the following dimensions: 1) the choice of the monetary aggregate, 2) the choice of the relevant variables to include in the X-vector, 3) the use of seasonally adjusted versus seasonally unadjusted data, 4) the specification of the time-series models used to describe expectations formation, 5) the sample period and, 6) the econometric estimation technique. Even though some of these model estimates should be more reliable than others for the reasons discussed earlier, the large number of estimates provide information on the sensitivity and reliability of the results reported here.

The results uniformly support the following conclusion. There is no empirical support here for the view that unanticipated increases in the money stock are negatively correlated with unanticipated changes in short interest rates. This conclusion is similar to that found in a previous paper, Mishkin (1981a), which conducts a parallel analysis of long-term interest rate behavior. However there are two aspects of the research methodology used here which raise questions about the general validity of this conclusion.

As has been discussed in the text, the β coefficients in the rational expectations models are not invariant to changes in the time-series processes of the money growth, income growth and inflation variables. Thus the conclusions derived from the estimates in this paper only provide information on the relationship of money growth and short rates for this postwar sample period. However, realize that many structural macroeconometric models which display a negative relationship between money growth and short rates have been estimated using a sample period which overlaps that used here. Thus the results reported in this paper are certainly of interest in evaluating these models.

A further difficulty with the research methodology followed here is that misspecification of (15), which describes expectations formation, could invalidate the results on the relationship between money growth and short rates. This is possible because misspecification of expectations formation could lead to inconsistent and biased β coefficients. However, the robustness of this paper's results to different specifications of the time-series models describing expectations provides evidence that this misspecification problem may not be very severe.

Given the conclusion reached above, how should we interpret it? If we are willing to accept exogeneity of the money supply process in the postwar period, the interpretation is clear cut. The evidence here would then cast doubt on the commonly held view that an unanticipated increase in the money stock will lead to an unanticipated decline in short-term interest rates. Not only does this suggest that the Federal Reserve cannot lower short interest rates by increasing the rate of money growth, but it also requires some modification of the monetary transmission mechanism embodied in structural macro-econometric models. It is plausible that an unanticipated increase in money growth may not induce unanticipated decline in short rates because it leads to an immediate upward revision in expected inflation. Thus, there is still a potential effect on real interest rates from unanticipated money growth and the evidence in no way denies that there are potent effects of money supply increases on aggregate demand.

As was mentioned in Section II of the paper, if unanticipated money growth is not exogeneous, then the β_m coefficient estimates are inconsistent and can lead to misleading inference. Particularly disturbing in this regard is the case where the Federal Reserve smooths interest rates so that an unanticipated increase in short rates causes a Federal Reserve reaction of an increase in unanticipated money growth. The resulting positive correlation of ε_t and $MG_t - MG_t^e$ would then tend to bias the β_m coefficient upward. Thus, even though the estimated β_m is positive, we cannot rule out the view in structural macroeconometric models that an exogenous increase in money growth leads to a decline in short rates, despite the empirical results of this paper.

Note however the nature of money growth endogeneity that is required for the above statement to be the case. If money growth is endogenous in the sense that the Federal Reserve modifies money growth within a quarter only in response to past public information available at the start of the quarter, this does not result in $MG_t - MG_t^e$ being correlated with ε_t. Hence the existence of Granger (1969) "causality" running from interest rates to money growth does not imply that the estimates of β_m will be inconsistent. Tests of the Sims (1972) variety therefore cannot shed light on the consistency of the β_m estimates. If we are not to reject the common view that increases in money growth lead to short interest rate declines, research of a fairly subtle sort is needed to demonstrate that unanticipated money growth is positively correlated with the contemporaneous error term, ε_t. Hence, this issue cannot be resolved without further research.

APPENDIX

Estimates of the Rational Expectations Models Using the Two-Step Procedure

The models in Table 3 were also estimated with the Barro (1977) two-step procefure over the 1959-76 sample period. The resulting coefficient estimates were not appreciably different from those in Table 3 with the M1G coefficients ranging from .20 to .29, the M2G coefficients from .09 to .16, the IPG coefficients from .03 to .06 and the π coefficients from .37 to .67. In order to gain further information on the robustness of the results, also estimated were rational expectations models which used residuals from eighth order autogressive models of the X-variables, as well as residuals from multivariate models of the X-variables which excluded the four lagged values of a variable only if they were not jointly significant at the ten percent level (rather than the five percent level as in the text). The results were quite close to those above, and again the evidence did not support a negative relationshop between unanticipated money growth and short rates.

Because the Federal Reserve may have changed its reaction function in the 1970's by paying more attention to the monetary aggregates than it did previously, it is possible that the results might change substantially if the 1970's are excluded from the sample period. Two-step estimates of the Table 3 models over the 1959-69 sample period did not support this conjecture. The money growth coefficients remained positive, although they did decline somewhat: the M1G coefficients ranged from .11 to .20, while the M2G coefficients ranged from .03 to .12. The IPG coefficients ranged from -.04 to .03 and the π coefficients from .37 to .53.

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