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5 Monetary Policy and Interest Rates: An Efficient Markets-Rational Expectations Approach

5.1 Introduction

The impact of a money stock increase on nominal interest rates is an important issue. 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 and medium runs, to a decline in interest rates. In these macro models (see Brainard and Cooper 1976; Modigliani 1974), the interest rate decline not only stimulates investment directly but also has a further expansionary impact on investment and consumer expenditure through its effect on the valuation of capital. The decline in interest rates is thus a critical element in the transmission mechanism of monetary policy. In addition, the view that increases in the money stock lead to an immediate decline in interest rates has important implications for the Federal Reserve System's conduct of monetary policy when a decline in interest rates is desired. This view is the basis for demands by government officials that the Fed not keep the rate of money growth too low and so induce an objectionable increase in interest rates.

Milton Friedman (1968, 1969) has criticized this view on the grounds that it ignores the dynamic effects of a money stock increase. Friedman concedes that a so-called liquidity effect—where an excess supply of money will create increased demand for securities, a rise in their price, and a resulting fall in interest rates—does work in the direction of a decline in interest rates when the money stock is increased. However, two other effects can counter this liquidity effect. The money stock increase will, over time, have an expansionary effect on both real income and the price level. This "income and price level effect" will, through the usual arguments in the money demand function, tend to reverse the decline in interest rates. More important for short-run effects on interest rates, increases in the money stock can also influence anticipations of inflation. Higher expected inflation as a result of money stock increases would, through a Fisher (1930) relation, increase nominal interest rates. This "price anticipations effect" can thus not only mitigate the decline in interest rates stemming from the liquidity effect but could also overpower it. That interest rates are highest in countries experiencing rapid rates of monetary growth is casual evidence for this proposition.

Early work on the issue of money supply increases and interest rates, such as Cagan (1972), Gibson (1970), and Gibson and Kaufman (1968), tended to stress the "income and price level effect" more than the "price anticipations effect" because these researchers believed that adjustments of inflationary expectations proceeded slowly. Recent work on the theory of rational expectations and market efficiency, starting with Muth (1961), indicates that inflationary expectations can adjust quite rapidly. Thus, the "price anticipations effect" should, and does, receive more weight in this chapter when the effect of money supply increases on interest rates is discussed.

Two lines of empirical work bear on the issue whether increases in the money stock lead to a decline or to a rise in interest rates. "Keynesian" macroeconometric models impose a fair amount of structure in their estimates of financial market and income-expenditure relationships. In these models, increases in the money stock do lead to a substantial decline in interest rates in the short and medium run, as, for example, in Modigliani (1974) and the simulation results in this chapter. This evidence is suspect, however, because these models ignore constraints that should be imposed if financial markets are efficient. Financial market efficiency cannot be ignored because evidence supporting it is quite strong (see Fama 1970). Furthermore, recent work (Mishkin 1978) indicates that a failure to impose financial market efficiency on macroeconometric models can yield highly misleading results.

An alternative empirical approach to this issue is to estimate reduced form relationships where changes in interest rates are regressed on past changes in the money stock. Evidence from this approach (Gibson and Kaufman 1968) does not support the view that increases in the money stock result in a fall in interest rates. Unfortunately, this evidence suffers from a problem endemic to reduced form empirical work: it is difficult to interpret the empirical results because the theoretical framework is obscure. Also, the absence of structure when changes in interest rates are regressed on changes in the money stock leads to a large number of parameters being estimated, and this results in statistical tests with low power.

Neither approach discussed above distinguishes between the effects from unanticipated versus anticipated monetary policy. Yet the theory of efficient capital markets and rational expectations does make this distinction, and this suggests an alternative approach to analyzing the relationship of money stock increases and interest rate movements. This chapter develops efficient-markets (or, equivalently, rational expectations) models for both long- and short-term interest rates and estimates them using postwar quarterly data. This approach has the advantage of imposing a theoretical structure on the problem that allows both easier interpretation of the empirical results and more powerful statistical tests of the proposition that increases in the money stock are correlated with declines in interest rates. Moreover, a Keynesian, liquidity preference view of interest rate determination can be embedded in the efficient-markets model and tested. Finally, as a side issue, attractive tests of bond market efficiency result from the approach used here.

5.2 The Model

The theory of efficient markets (or, equivalently, rational expectations) implies that interest rates in a bond market should reflect all available information. More precisely, it implies that the market uses available information correctly in assessing the probability distribution of all future interest rates and bond returns. Hence for long bond returns, y_t, and short-term interest rates, r_t,

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

and

$$(2) \qquad E_m(r_t|\phi_{t-1}) = E(r_t|\phi_{t-1}),$$

where

y_t = the one-period (from $t - 1$ to t) nominal return from holding long-term bonds—it includes capital gains plus interest payments,
r_t = the one-period (short-term) 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 expectation assessed by the market at $t - 1$.

In order to give this concept empirical content we must specify models of market equilibrium. For the case of long-term bonds, we assume, as in the previous chapter, that the market equates expected one-period holding returns across securities, allowing for risk (liquidity) premiums which are constant over time. This model of market equilibrium implies that

$$(3) \qquad E_m(y_t|\phi_{t-1}) = r_{t-1} + d^l,$$

where d^l = a constant liquidity premium for long-term bonds.

A more refined model of market equilibrium allowing the risk premium to vary over time is not used for long-term bonds because, as discussed in the previous chapter, it makes little difference to the empirical tests. Combining the model of market equilibrium above with market efficiency yields the same condition as in Chapter 4:

(4) $$E(y_t - r_{t-1} - d^l|\phi_{t-1}) = 0$$

and the same efficient-markets model

(5) $$y_t = r_{t-1} + d^l + (X_t - X_t^e)\beta^l + \epsilon_t^l,$$

where an e superscript denotes expected values on all past information [i.e., a rational forecast is defined as $X_t^e = E(X_t|\phi_{t-1})$], and

X_t = a variable (or vector of variables) relevant to the pricing of bonds,
β^l = a coefficient (or vector of coefficients),
ϵ_t^l = serially uncorrelated error process [because $E(\epsilon_t^l|\phi_{t-1}) = 0$].

In the analysis of short-term interest rates, we can no longer argue that the model of the market equilibrium has no effect on tests of the efficient-markets model. In this case, the model of the risk premium used here does contribute significantly to the explanation of the dependent variable. We assume, as in Fama (1976*b*), that the one-period-ahead forward rate equals the one-period-ahead expected short rate, plus a risk premium that now varies over time with the uncertainty in short-rate movements, that is,

(6) $${}_{t-1}F_t = E_m(r_t|\phi_{t-1}) + d_t^s$$

and

(7) $$d_t^s = a_0 + a_1\,\sigma_t,$$

where

${}_{t-1}F_t$ = forward rate for the one-period rate at time t implied by the yield curve at $t-1$,
d_t^s = risk (liquidity) premium for ${}_{t-1}F_t$,
σ_t = measure of uncertainty in short rate movements.

Combining this model of market equilibrium with the rationality or market efficiency condition of (2) yields

(8) $$E(r_t - {}_{t-1}F_t - a_0 - a_1\,\sigma_t|\phi_{t-1}) = 0$$

and the corresponding efficient-markets model

(9) $$r_t = {}_{t-1}F_t - a_0 - a_1\sigma_t + (X_t - X_t^e)\beta^s + \epsilon_t^s,$$

where the s superscript is used to differentiate the β and ϵ from their counterparts in the long-term bond model.

The research question posed in the first section of this chapter suggests that money growth is an interesting piece of information relevant to the pricing of bonds and interest rates. Substituting for X_t leads to the following efficient-markets models:

$$(10) \qquad y_t = r_{t-1} + d^l + \beta^l_m(MG_t - MG^e_t) + \epsilon^l_t,$$

$$(11) \qquad r_t = {}_{t-1}F_t - a_0 - a_1\sigma_t + \beta^s_m(MG_t - MG^e_t) + \epsilon^s_t,$$

where MG_t = the money growth rate at time t.

As is found in the foreign exchange market (see Mussa 1979), spot and forward rates move together, so that changes in short-term interest rates are predominantly unanticipated. Because the long rate is closely linked to the price of a long bond, over periods as short as a quarter the correlation of changes in long rates with unanticipated bond returns, $y_t - r_{t-1} - d$, is very negative: $-.96$ in the sample period used in the following empirical work. Changes in long interest rates will thus also be predominantly unanticipated. We can see how the efficient-markets models above differ from earlier analysis: they stress that only unanticipated movements in money growth can have an effect on unanticipated movements in interest rates. Since changes in interest rates are predominantly unanticipated, these efficient-markets models emphasize the effects of unanticipated money growth movements on changes in interest rates. In contrast, the earlier work does not distinguish between the effects of anticipated and unanticipated money growth.

If unanticipated increases in money growth are associated with a decline in long rates (as might be expected from "Keynesian" macroeconometric models), the coefficient on unanticipated money growth should be significantly positive in the long bond equation above because $y_t - r_{t-1}$ and the change in long rates are negatively correlated: that is, $\beta^l_m > 0$. If unanticipated increases in money growth are associated with a decline in short rates, then the coefficient on unanticipated money growth in the short-rate equation should be significantly negative, that is, $\beta^s_m < 0$.

An important caveat is in order. As noted in Chapter 2, the efficient-markets model does not guarantee that $X_t - X^e_t$ is exogenous so that the estimates of β are consistent. Another way to make this point is to acknowledge that the efficient-markets model does not indicate whether a significant β coefficient implies causation from its unanticipated variable to bond prices and interest rates. As far as market efficiency is concerned, causation could just as well run in the other direction, or it could be nonexistent, as in the case where new information simultaneously affects both interest 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 same caveat applies especially when we analyze the estimated β_m coefficient. If the money supply process is seen as exogenous, the interpretation of the estimated β_m's is straightforward. The finding of a significant positive β^l_m and negative β^s_m will then provide evidence supporting the "Keynesian" position that increased money growth leads, at least in the short run, to declines in interest rates; and a failure to find this result will cast doubt on this view. If the money supply process is not exogenous, however—the position taken by many critics of monetarist analysis—then the estimated β_m coefficients may suffer from simultaneous equation bias and give a misleading impression about how increases in the money supply affect interest rates. Because this chapter provides no evidence on the exogeneity of the money supply process, the β_m estimates must be viewed as providing information only on the correlations of unanticipated money growth and the movements in interest rates. Interpretation of these correlations is deferred to the concluding remarks at the end of the chapter.

The liquidity preference approach to the demand for money (see Goldfeld 1973; and Laidler 1977) indicates that interest rates are related not only to money growth but also to movements in real income, the price level, and inflation. Adding this information to the X vector of the efficient-markets models, noting that unanticipated inflation is equivalent to the unanticipated price level, leads to the following:

$$y_t = r_{t-1} + d^l + \beta^l_m(MG_t - MG^e_t) + \beta^l_y(YG_t - YG^e_t) + \beta^l_\pi(\pi_t - \pi^e_t) + \epsilon^l_t \tag{12}$$

$$r_t = {}_{t-1}F_t - a_0 - a_1\sigma_t + \beta^s_m(MG_t - MG^e_t) + \beta^s_y(YG_t - YG^e_t) + \beta^s_\pi(\pi_t - \pi^e_t) + \epsilon^s_t \tag{13}$$

where

YG_t = growth rate of real income,
π = inflation rate,
$\beta_m, \beta_y, \beta_\pi$ = coefficients.

These equations are really efficient-markets analogs to the typical money demand relationship found in the literature. In addition, they capture elements of interest rate models of the Feldstein and Eckstein (1970) variety.

The magnitude and sign of the β coefficients in equations (10)–(13) depend on the time-series processes of the money supply, real income, and price level, even when the sign and magnitude of these coefficients are assumed to reflect an underlying structural theory such as liquidity preference. If the time-series processes of real income and the price level are such that an unanticipated rise in these variables is not followed by more than a compensating decline in these variables, then a liquidity preference view implies that the coefficients of unanticipated income growth and inflation should be negative in the long bond equation (12)—that is, $\beta^l_y < 0$, $\beta^l_\pi < 0$—and positive in the short-rate equation (13)—that is, $\beta^s_y > 0$, $\beta^s_\pi > 0$. In this case, an unanticipated increase in income growth should lead to higher interest rates, currently and in the future. The negative effect of an unanticipated increase in inflation on bond returns follows from the resulting reduction in real money balances, which also leads to rising interest rates. The unanticipated inflation effect should be further strengthened if, as in the Cagan (1956) adaptive expectations model, expected inflation rises when actual inflation is above its expected value. In this case, an unanticipated rise in inflation promotes a rise in nominal interest rates either through a Fisher (1930) relation or because expected inflation is a separate argument in the money demand function, as in Friedman (1956).

Note also that the more persistent the time-series process of inflation and income growth—that is, the more an unanticipated increase in these variables leads to further increases—the more powerful the unanticipated income and inflation effects on interest rates indicated by the theoretical structure discussed above. Clearly, the importance of the "income and price level" and "price anticipations" effects also depend on the time-series process of money growth. Thus the β_m coefficients also will not be invariant to changes in the money growth, time-series process.

We now turn to the actual estimation of the efficient-markets models of equations (10)–(13), with the warning that some caution must be exercised when interpreting results from estimates of these equations because the direction of causation cannot be established in this framework.

5.3 Empirical Results

5.3.1 The Data

Postwar quarterly data is used in the empirical work below. The long bond models are estimated over the sample period 1954–1976. However, six-month Treasury Bills were not issued before 1959, and since the six-month bill rate is needed to calculate the forward rate in the short rate models, these models are estimated over the 1959–1976 period. The data sources and definitions of variables used in these estimations are as follows:

y_t = quarterly return from holding a long-term government bond from the beginning to the end of the quarter,

r_t = the ninety-day Treasury Bill rate, the last trading day of the quarter in fractions at an annual rate in the short rate equations, but r_{t-1} is at a quarterly rate in the long bond equations,

$$_{t-1}F_t = \quad 4\left[1 - \frac{(360 - 180\, rsix_{t-1})}{(360 - 90 r_{t-1})}\right],$$

where

$rsix_t$ = the six-month (180 days) bill rate at the end of quarter—in fractions at an annual rate,

$M1G_t$ = 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,

$M2G$ = 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.

Unless otherwise noted, all these variables have been constructed from seasonally adjusted data except for r_t, ${}_{t-1}F_t$, and y_t, which do not require seasonal adjustments. The bond return series was obtained from the Center for Research in Security Prices (CRSP) at the University of Chicago and is described in Fisher and Lorie (1977) and Mishkin (1978). The Treasury Bill data was supplied by the Federal Reserve Board. The IPG_t and π_t variables were constructed from data in the Department of Commerce's *Business Statistics* and *Survey of Current Business*. 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*. All other variables used to specify the forecasting equations were obtained from the NBER data bank.

As shown in Chapter 4, using averaged data in efficient-markets or rational expectations models can give misleading results. The data for bond returns and interest rates here are thus derived from security prices at particular points in time. For the same reason, the derivation of the other variables here uses data as close to being end of quarter as possible. Industrial Production is thus made a proxy for real income in estimating the models rather than the more broadly based national income accounts measure. Similarly, the CPI has been used to calculate the inflation variable rather than the GNP deflator. Endpoint data (or close to endpoint) help unearth significant relationships between bond returns and unanticipated variables. Some experiments with quarterly averaged data led to worse fits for efficient-markets models, fewer significant coefficients, and no appreciable differences as to the statistical significance of the β_m coefficients.

5.3.2 The Estimation Procedure

To estimate the short and long rate models of equations (10)–(13), measures of anticipated money growth, industrial production growth and inflation are needed. Here, these anticipations are assumed to be rational forecasts obtained from linear forecasting equations. The model estimates are produced by estimating each short or long rate equation jointly with the forecasting equations, and imposing the cross-equation restrictions implied by rationality of expectations. See Chapter 2 for details of this procedure.

In Chapter 2 we saw that economic theory may not be a useful tool for deciding on the specification of the forecasting equations. Thus atheoretical statistical procedures are used here. If indeed economic theory is not particularly useful in evaluating the forecasting equations, it is all the more important to check for the robustness of results to changes in the specification of these equations. Therefore, two procedures for specifying the forecasting equations are used in the text, and results with several additional specifications are discussed in Appendix 5.2.

The simplest forecasting equations are univariate time-series models of the autoregressive type. Fourth-order autoregressions are usually successful in reducing residuals in quarterly data to white noise and are thus used here. Ordinary least-squares estimates for the $M1G$, $M2G$, IPG, and π equations for both sample periods used in estimation can be found in Appendix 5.1. There is a fair amount of persistence in the time-series models for money growth and inflation, indicating that "income and price level" and "price anticipation" effects of the sort that Friedman (1968, 1969) discusses are potentially important. Although less persistence is evident in the time-series process of industrial production growth, it has the characteristic that a positive innovation does lead to a permanently higher level of industrial production (although not in the rate of growth). Thus, as discussed in the preceding section, the unanticipated inflation and IPG coefficients may be expected to be negative in the long-rate bond equations and positive in the short-rate equations.

The univariate time-series models suffer from the problem of unstable coefficients. Chow (1960) tests reported in Appendix 5.1, where the sample period has been split into equal lengths, reject in five out of eight equations the hypothesis that the coefficients of the univariate models are equal in the two subperiods. Multivariate forecasting equations thus have been specified by the procedure outlined in Chapter 2 which makes use of Granger's (1969) concept of predictive content. 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 ninety-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. (These other variables were selected because a reading of the literature on Federal Reserve reaction functions—see Fair 1978 and the references therein—indicated that they might help explain money growth.) The four lagged values of each variable were retained in the equation only if they were jointly significant at the 5 percent level or higher.

The resulting multivariate time-series models for both sample periods can be found in Appendix 5.1 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. Not only do these models have a better fit than the corresponding univariate models, but Chow tests reported in Appendix 5.1 now reject stability of the coefficients in only one out of eight cases.

Before we turn to the empirical results, the measure of short-rate uncertainty, σ_t, used here requires some discussion. Fama (1976*b*) calculates σ_t as the average of the absolute value 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 $t-1$ could be problematic. An alternative, though similar, measure of σ_t is used in this study. The difference between the forward rate and the spot rate, that is, $r_t - {}_{t-1}F_t$, 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:

$$r_t - {}_{t-1}F_t = \underset{(.0017)}{-\ .0001} \ \underset{(.2937)}{-\ 1.0961}\ \sigma_t + \epsilon_t \tag{14}$$

$R^2 = .1659$ standard error $= .0068$
Durbin-Watson $= 1.90$

where

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

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

5.3.3 The Results

There is no strong theoretical reason to estimate the long bond or short-rate model with one monetary aggregate versus another. Unanticipated growth rates of both $M1$ and $M2$ are therefore used and results with additional monetary aggregates are explored in Appendix 5.2. The resulting estimates of the models appear in tables 5.1 and 5.2. Panel A of

Table 5.1 **Nonlinear Estimates of the Long Bond Model Using Seasonally Adjusted Data**

	Coefficients of				
Model	$M1G-M1G^e$	$M2G-M2G^e$	$IPG-IPG^e$	$\pi-\pi^e$	Constant Term
	Panel A: Using Univariate Forecasting Models				
1.1	.0482				−.0014
	(.5961)				(.0032)
1.2	.0501		−.4242**	−1.8482*	−.0029
	(.5517)		(.1260)	(.8028)	(.0032)
1.3		.9174			−.0013
		(.5459)			(.0032)
1.4		.7174	−.4077**	−1.7691*	−.0027
		(.5063)	(.1243)	(.7880)	(.0031)
	Panel B: Using Multivariate Forecasting Models				
1.5	−.2621				−.0014
	(.7429)				(.0032)
1.6	.4108		−.5039**	−1.7529	−.0020
	(.7164)		(.1568)	(.9552)	(.0032)
1.7		.9199			−.0014
		(.6738)			(.0032)
1.8		1.0950	−.4987**	−1.8206	−.0020
		(.6283)	(.1492)	(.9353)	(.0032)

Note: Asymptotic standard errors of coefficients in parentheses.
* = Significant at the 5 percent level.
** = Significant at the 1 percent level.

these tables contains estimates which make use of the univariate forecasting equations, while panel B's estimates use the multivariate forecasting models of the form found in Appendix 5.1. The estimates of the γ coefficients are not presented here because they are not particularly interesting.

An issue basic to these results is whether the efficient-markets (rational expectations) model used here is valid. Previous evidence on the efficiency of the bond market indicates that efficient-markets models of the type used here are a reasonable characterization of bond market behavior. Table 5.3 contains likelihood ratio tests, described in Chapter 2, of the nonlinear constraints implied by both market efficiency (rational expectations) and the model of market equilibrium. The test statistics do not reject the constraints for any of the long bond models in table 5.1: the marginal significance level of the statistics are never lower than .05. These results then also provide additional evidence for the efficient-markets model of long bond behavior.

Table 5.2 **Nonlinear Estimates of the Short Rate Model Using Seasonally Adjusted Data**

	Coefficients of					
Model	$M1G - M1G^e$	$M2G - M2G^e$	$IPG - IPG^e$	$\pi - \pi^e$	Constant Term	σ
	Panel A: Using Univariate Forecasting Models					
2.1	.2788*				.0006	−1.2266**
	(.1088)				(.0015)	(.2714)
2.2	.2774**		.0352	.6211**	.0002	−1.1514**
	(.1075)		(.0275)	(.1989)	(.0014)	(.2618)
2.3		.1616			.0006	−1.2563**
		(.1085)			(.0015)	(.2851)
2.4		.1904	.0399	.6545**	.0002	−1.1571**
		(.1053)	(.0278)	(.2058)	(.0015)	(.2686)
	Panel B: Using Multivariate Forecasting Models					
2.5	.1677				.0006	−1.2761**
	(.1283)				(.0015)	(.2863)
2.6	.2512		−.0455	.5199	.0004	−1.2109**
	(.1381)		(.0493)	(.3272)	(.0016)	(.3015)
2.7		.2562			.0001	−1.1807**
		(.1341)			(.0016)	(.2917)
2.8		.3039*	−.0770	.6501*	−.0004	−1.0779**
		(.1409)	(.0471)	(.3314)	(.0016)	(.3069)

Note: See table 5.1.

The table 5.3 results for the short-rate models, however, reject the nonlinear constraints at the 5 percent level in six out of eight cases. How should we interpret these rejections? They can result from either the failure of rationality (market efficiency) or of the model of market equilibrium. Both models of market equilibrium used in the long bond and short-rate models are crude: neither risk premium is derived from utility maximizing behavior. A theoretically more justifiable risk premium would, for example, exploit the covariance of bill or bond returns with returns on alternative assets. Yet, as the regression results in equation (14) indicate, the model of market equilibrium is a significant element in explaining the movements of the dependent variable in the short-rate equation. In this situation, unlike that for the long-rate equation where the model of market equilibrium appears to be unimportant in explaining the dependent variable, its misspecification can lead to rejections of the nonlinear constraints. Thus, rejections of the nonlinear constraints occurring in the short-rate models, but not in the long bond models, can be attributed plausibly to misspecification in the model of market equilibrium.

Table 5.3 **Likelihood Ratio Tests of Nonlinear Constraints**

Model	Likelihood Ratio Statistic	Marginal Significance Level
1.1	$\chi^2(4) = 6.45$	.1680
1.2	$\chi^2(12) = 14.00$	.3007
1.3	$\chi^2(4) = 3.20$	.5249
1.4	$\chi^2(12) = 12.43$	.4118
1.5	$\chi^2(12) = 18.10$	.1127
1.6	$\chi^2(24) = 33.81$	.0881
1.7	$\chi^2(8) = 10.67$	.2211
1.8	$\chi^2(24) = 32.60$	.1128
2.1	$\chi^2(4) = 12.76$*	.0125
2.2	$\chi^2(12) = 13.65$	.3235
2.3	$\chi^2(4) = 12.46$*	.0143
2.4	$\chi^2(12) = 17.18$	.1430
2.5	$\chi^2(8) = 21.65$**	.0056
2.6	$\chi^2(28) = 50.02$**	.0067
2.7	$\chi^2(8) = 25.69$**	.0012
2.8	$\chi^2(28) = 50.92$**	.0051

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

*Significant at the 5 percent level.

**Significant at the 1 percent level.

A suitable transformation of the unconstrained system outlined in Chapter 2 yields additional evidence on the potential misspecification of the model of market equilibrium. The unconstrained system where the γ are not equal in the forecasting and short-rate equations can be rewritten as

$$(15) \qquad \begin{aligned} X_t &= Z_{t-1}\gamma + u_t, \\ r_t &= {}_{t-1}F_t - a_0 - \sigma a_1 - Z_{t-1}\alpha + (X_t - Z_{t-1}\gamma)\beta^s + \epsilon_t, \end{aligned}$$

where the γ's are constrained to be equal in the two equations. Therefore, the nonlinear 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 risk premium is related to the variables in Z, yet they have been excluded from the model of the risk premium, then this may explain the rejections of the nonlinear constraints. To make this conjecture plausible, we should expect that a model of the risk premium which is related to Z would have reasonable characteristics. The Fama-type model, for example, does generate plausible values. The resulting risk premiums (at annual rates) have a mean of 57 basis points (1/100 of a percentage point) and a standard deviation of 30 basis points. They also move smoothly: their autocorrelations for lags of one through four are respectively .96, .91, .85, and .78. In the model which leads to the strongest rejection of the nonlinear constraints, model 2.7, we could attribute this rejection to the fact that a more appropriate specification of the risk premium is $d_t^s = a_o + a_1\sigma_t + Z_{t-1}\alpha$, where Z_{t-1} contains the four lagged values of money growth ($M2G$) and Treasury Bill rates (r). This latter specification leads to values for the risk premium that are somewhat more variable and less smooth than the equation (14) specification, but not appreciably so. The risk premium from this expanded specification has 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 (15) also has the advantage of providing us with potentially interesting information on the risk premium. The results indicate that the 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 5.A.2 of Appendix 5.1. The results here thus point out a direction for future research on the risk premium. Following Nelson (1972), I also conducted more direct experiments on the relation of the risk premium to lagged interest rates and unemployment with negative results. Experiments with lagged values of $r-F$ also did not add explanatory power to the model of the risk premium.

If a misspecified model of the risk premium is the source of the rejections of the nonlinear constraints, the efficient markets-rational expectations model used here is fortunately still a valid framework for analyzing the relationship of money growth and short rates. With rational expectations, the unanticipated $X_t - X_t^e$ variable 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, they will be orthogonal to $X_t - X_t^e$. The exclusion of these variables, and the resulting rejection of the nonlinear constraints, will not lead to inconsistent estimates of the β coefficients. Since it is not necessary to derive a better model of the risk premium to achieve reliable estimates of the β's, this tricky research issue, which is beyond the scope of this study, is left to future research.

The unanticipated $M1G$ coefficients in table 5.1 do not lend support to the view that an unanticipated increase in money growth is correlated with a fall in long bond rates. In panel A, although both of these coefficients have a positive sign, they are not significantly different from zero at the 5 percent level: asymptotic t statistics are less than .1. In addition these coefficients are quite small. The β coefficients here denote the percentage point change in the bond return from a 1 percent error in anticipations, and in our 1954–1976 sample period, a one percentage point rise in the quarterly bond return corresponds approximately to a 10 basis point (1/100 of a percentage point) fall in the long bond rate. Thus, the $M1G$ coefficients in panel A indicate that a 1 percent surprise increase in $M1$ is associated with only a .5 basis point decline in the long bond rate.

The panel B estimates of the $M1G$ coefficients indicate that the conclusion on the relationship of long rates and money growth is not altered by using multivariate versus univariate forecasting models in estimation. Again, neither of the unanticipated $M1G$ coefficients are significantly different from zero at 5 percent, and they continue to be small, with the largest of the coefficients indicating that a 1 percent surprise increase in $M1$ leads to only a 4.1 basis point decline in the long bond rate. Furthermore, one of the unanticipated $M1G$ coefficients is now negative.

The coefficients on unanticipated $M2$ growth in table 5.1 are more positive than the unanticipated $M1G$ coefficients, they nevertheless do not lend strong support to the view that unanticipated money growth should be negatively correlated with the change in long rates. They do not differ significantly from zero at the 5 percent level (although in 1.3 the unanticipated $M2G$ coefficient is significantly different from zero at the 10 percent level). Also note that the $M2$ results in panel A and in panel B are so similar that it is again clear that the results on unanticipated money growth are not particularly sensitive to specifications of the money growth forecasting model.

How different are these findings from those that might be inferred from "Keynesian," structural macroeconometric models? Using a simulation technique discussed in Mishkin (1979) we can examine the response of a macromodel to a 1 percent surprise increase in $M1$. Equation 1.1 was used to trace out the effect on $M1$ growth from a 1 percent innovation. The resulting $M1$ numbers were then used in a simulation experiment with the MPS (MIT-Penn-SSRC 1977) Quarterly Econometric Model in order to derive the response of this model to the 1 percent $M1$ innovation occurring in the 1967:1 quarter. The MPS model indicates that this 1 percent $M1$ innovation would lead to an immediate decline of 18.1 basis points in the long rate. Not only is this long-rate decline several times larger than the maximum 4.1 basis point decline implied by the empirical evidence in table 5.1, but also it is significantly larger at the 5 percent level for three of the four estimates in table 5.1 (and is almost significantly larger for the remaining estimate). Clearly, the coefficients on unanticipated $M1$ growth are quite low relative to what might be expected from a structural macroeconometric model.

The unanticipated inflation and industrial production coefficients in table 5.1 conform to our priors. In both the $M1$ and $M2$ efficient-market models, these coefficients are negative and are either significant or nearly significant at the 5 percent level. The results on the coefficients of unanticipated industrial production growth are especially strong, with both the panel A and panel B estimates significantly different from zero at the 1 percent level. Although the unanticipated inflation coefficients are very similar in both panels, their asymptotic standard errors rise somewhat from panel A to panel B. They are thus not quite significant at the 5 percent level in panel B, while they are significant at this level in panel A.

The similarity between the money growth as well as other coefficient estimates in panel A and panel B is encouraging, for it gives us confidence that these results are robust to changes in the models describing expectations. Further model estimates described in Appendix 5.2 with additional specifications for the forecasting equations yield similar results. This is an important finding. Poor specification of expectations formation appears to be a major concern in this line of research because it leads to errors-in-variables bias in the coefficient estimates. The important question is, How severe would this bias be? Denoting the measured $X_t - X_t^e$ by an m superscript and the true $X_t - X_t^e$ with a T superscript, we can write

$$(16) \qquad (X_t - X_t^e)^m = (X_t - X_t^e)^T + v_t,$$

where v_t is the measurement error. If such variables as money growth, industrial production growth, and inflation are hard to forecast—which seems likely—then the variance of the true $X_t - X_t^e$ forecast error will be substantial. If the incremental predictive power of other information besides the past history of the forecasted variable is not large, then the variance of the measurement error in expectations used here will be small in relation to the variance of the true forecast error: that is, Var $[(X_t - X_t^e)^T] >> \text{Var}(v_t)$. If this occurs, the errors-in-variables bias would be negligible and should not be an important problem in this research.

The similarity of the model estimates despite substantial changes in the specifications for the forecasting equations is found not only in this chapter, but also in the chapters preceding and following. This provides strong support for the view that unanticipated increases in money growth are associated with interest rate declines. Moreover, the smaller standard errors found for the coefficients estimated using univariate rather than multivariate forecasting equations provides some support for the position taken by Feige and Pearce (1976), that forecasts from univariate time-series models may be "economically rational" expectations.[1]

The results for the short-rate model in Table 5.2 are even more damaging to the view that associates a decline in interest rates with an unanticipated money growth increase.[2] All the coefficients on both unanticipated $M1$ and $M2$ growth are positive in table 5.5, and in three cases the coefficients are statistically significant. They indicate that a 1 percent

1. Note that Sims (1977) has raised some questions about the statistical techniques used by Feige and Pearce (1976), and this does cast some doubt on their evidence.

2. Urich and Wachtel (1981) obtain similar results using weekly data. Thus, reduction of the unit of observation in the analysis is likely to leave the findings here intact.

Table 5.4 **Nonlinear Estimates of the Long Bond Model Using Seasonally Unadjusted Data**

Model	$M1G-M1G^e$	$M2G-M2G^e$	$IPG-IPG^e$	$\pi-\pi^e$	Constant Term
	Coefficients of				
	Panel A: Using Univariate Forecasting Models				
4.1	−.7339*				−.0017
	(.3631)				(.0031)
4.2	−.5879		−.2028*	−2.5145**	−.0026
	(.3631)		(.0857)	(.6912)	(.0032)
4.3		.0001			−.0014
		(.3610)			(.0032)
4.4		.1426	−.2420**	−2.4438**	−.0024
		(.3330)	(.0838)	(.7111)	(.0032)
	Panel B: Using Multivariate Forecasting Models				
4.5	−1.2781**				−.0014
	(.4504)				(.0032)
4.6	−.8078		−.4105**	−2.4472**	−.0020
	(.4339)		(.1371)	(.8089)	(.0032)
4.7		−.1404			−.0014
		(.4821)			(.0032)
4.8		.1534	−.4741**	−2.6226**	−.0020
		(.4391)	(.1396)	(.8237)	(.0033)

Note: See table 5.1.

surprise increase in $M1$ or $M2$ is associated with a 16–30 basis point unanticipated increase in the bill rate. The simulation experiment with the MPS model that is described above indicates that a 1 percent $M1$ surprise leads to an immediate *decline* of 88 basis points in the bill rate. This finding contrasts strongly with the finding here that even the least positive $M1$ coefficient is more than eight of its standard errors away from this figure.

The results on the unanticipated inflation coefficients are similar to those in the long bond model. These coefficients are positive, as might be expected, and are significantly different from zero in three out of four cases. The results on the unanticipated industrial production growth coefficients are not quite as strong as in the earlier table. They are never statistically significant, and in panel B they even have the wrong sign.

The efficient markets–rational expectations model does not specify whether the $X - X^e$ variables should be described by seasonally adjusted or seasonally unadjusted data. This empirical issue cannot be settled easily on theoretical grounds because it is not clear whether market participants concentrate on seasonally adjusted versus unadjusted in-

Table 5.5 **Nonlinear Estimates of the Short Rate Model Using Seasonally Unadjusted Data**

	Coefficients of					
Model	$M1G-M1G^e$	$M2G-M2G^e$	$IPG-IPG^e$	$\pi-\pi^e$	Constant Term	σ
	Panel A: Using Univariate Forecasting Models					
5.1	.3029**				.0003	−1.1255**
	(.0652)				(.0014)	(.2530)
5.2	.2458**		.0274	.4687**	.0001	−1.1267**
	(.0671)		(.0171)	(.1716)	(.0014)	(.2464)
5.3		.1926*			.0003	−1.1468**
		(.0644)			(.0015)	(.2765)
5.4		.1967**	.0440*	.5459**	.0001	−1.1260**
		(.0624)	(.0176)	(.1746)	(.0014)	(.2526)
	Panel B: Using Multivariate Forecasting Models					
5.5	.3431**				.0007	−1.2403**
	(.0831)				(.0015)	(.2639)
5.6	.2484**		.0386	.5079	.0004	−1.2037**
	(.0956)		(.0376)	(.2623)	(.0016)	(.2861)
5.7		.3285**			−.0007	−.9891**
		(.0918)			(.0015)	(.2841)
5.8		.2011*	.0400	.5788*	−.0003	−1.0791**
		(.0986)	(.0374)	(.2674)	(.0016)	(.3021)

Note: See table 5.1.

formation. For this reason, the table 5.1 and table 5.2 models have also been estimated with seasonally unadjusted data for the X's. The resulting estimates and test statistics appear in tables 5.4, 5.5, and 5.6 and were obtained by the same procedures as the previous estimates with seasonally adjusted data.

A comparison of tables 5.4–5.6 with tables 5.1–5.3 indicates that the choice of adjusted or unadjusted data is not a critical factor in this research. The likelihood ratio tests of the nonlinear constraints yield similar conclusions. In addition the coefficient estimates are similar, although their standard errors tend to be smaller in the seasonally unadjusted results. In the short-rate models, all the industrial production growth coefficients now have the "correct" positive sign.

The seasonally unadjusted data are even less favorable to the view that increased money growth is associated with a decline in interest rates. Now all the $M1$ coefficients in the long bond model are negative, implying a positive correlation of movements in money growth and long interest rates, and two of these coefficients are significantly different from zero at the 5 percent level. In addition, the $M2$ coefficients in the long-rate model

Table 5.6 **Likelihood Ratio Tests of Nonlinear Constraints**

Model	Likelihood Ratio Statistic	Marginal Significance Level
4.1	$\chi^2(4) = 5.08$	.2792
4.2	$\chi^2(12) = 20.83$	.0529
4.3	$\chi^2(4) = 3.27$	.5137
4.4	$\chi^2(12) = 19.53$	.0765
4.5	$\chi^2(12) = 12.10$	.4377
4.6	$\chi^2(24) = 28.48$	.2403
4.7	$\chi^2(8) = 7.23$	.5120
4.8	$\chi^2(24) = 27.11$	.2994
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: See table 5.3.

are less positive. For the short-rate models, all the money growth coefficients in table 5.4 are positive and are now statistically significant at the 5 percent level, with six out of eight significant at the 1 percent level. The seasonally unadjusted data, then, lend support to the contrary view that unanticipated movements in money growth and interest rates are positively correlated.

5.4 Concluding Remarks

A wide range of empirical tests of the relationship between money growth and interest rates have been conducted in this chapter and in Appendix 5.2. A guiding principle of this research has been to use many different empirical tests of the model in order to provide information on the robustness of the results. In pursuit of this goal, model estimations have been varied 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 forecasting models used to describe expectations formation, and (5) the sample period. The large number of estimates provide information on the sensitivity and reliability of the results reported here.

The results point to the following conclusions. There is no empirical support for the view that an unanticipated increase in the money stock is associated with a decline in interest rates. However, there are two aspects of the research methodology used here which raise questions about the general validity of this conclusion.

As we have seen, the β coefficients in the efficient markets–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 from the estimates in this chapter provide information on the relationship between money growth and interest rates only for the postwar sample period. However, realize that many structural macroeconometric models displaying a negative relationship between money growth and interest rates have been estimated using sample periods which overlap those used here. The results reported in this chapter are certainly of interest in evaluating these models.

A further difficulty with the present research methodology is that misspecification of the forecasting equations describing expectations formation can invalidate the results on the relationship between money growth and interest rates. Specifically, misspecification of expectations formation can lead to inconsistent and biased β coefficients. However, the robustness of results to different specifications of the time-series models describing expectations provides evidence that the misspecification problem is probably not very severe.

How should we interpret the conclusion reached above? If we are willing to accept exogeneity of the money supply process in the postwar period and the efficient-markets models as true reduced forms, 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 a decline in interest rates. Not only does this suggest that the Federal Reserve cannot lower interest rates by increasing the rate of money growth, but it also requires some modification of the monetary transmission mechanism embodied in structural macroeconometric models. It is plausible that an unanticipated increase in money growth will not induce a decline in interest 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 5.2, 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 is interest rates causes the Federal Reserve to react by an unanticipated increase in money growth. The resulting correlation of $MG_t - MG_t^e$ with the ϵ_t error terms (negative with ϵ_t^l and positive with ϵ_t^s) tends to bias the results toward a positive association of money growth and interest rates. Thus, we cannot rule out the view in structural macroeconometric models that an exogenous increase in money growth leads to a decline in interest rates, despite the empirical results of this chapter.

Note, however, the nature of money growth endogeneity required for this reservation to hold. 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, $MG_t - MG_t^e$ will not be correlated with ϵ_t^l or ϵ_t^s. 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. "Causality" 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 interest rate declines, research of a fairly subtle sort is needed to demonstrate that unanticipated money growth is negatively correlated with ϵ_t^l and positively correlated with ϵ_t^s. Hence, this issue cannot be resolved without further research.

Appendix 5.1: Estimates of the Forecasting Equations

Table 5.A.1 Univariate Time-Series Models

Model:	A1.1	A1.2	A1.3	A1.4	A1.5	A1.6	A1.7	A1.8
Dependent Variable:	*M1G*	*M1G*	*M2G*	*M2G*	*IPG*	*IPG*	π	π
Sample Period:	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976
Constant term	.0038	.0053	.0049	.0076	.0093	.0092	.0012	.0014
	(.0013)	(.0019)	(.0017)	(.0024)	(.0283)	(.0035)	(.0066)	(.0008)
$M1G(-1)$	.3777	.3158						
	(.1078)	(.1234)						
$M1G(-2)$	.2205	.2166						
	(.1152)	(.1289)						
$M1G(-3)$	.0550	.0306						
	(.1151)	(.1275)						
$M1G(-4)$	−.0341	−.0371						
	(.1076)	(.1236)						
$M2G(-1)$			.6598	.6113				
			(.1080)	(.1220)				
$M2G(-2)$			−.0542	−.0412				
			(.1280)	(.1421)				
$M2G(-3)$			.1736	.1619				
			(.1286)	(.1408)				
$M2G(-4)$			−.0758	−.1525				
			(.1094)	(.1230)				

Table 5.A.1 (continued)

Model:	A1.1	A1.2	A1.3	A1.4	A1.5	A1.6	A1.7	A1.8
Dependent Variable:	$M1G$	$M1G$	$M2G$	$M2G$	IPG	IPG	π	π
Sample Period:	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976
$IPG(-1)$					.4254	.3514		
					(.1003)	(.1187)		
$IPG(-2)$					−.2346	−.2100		
					(.1091)	(.1250)		
$IPG(-3)$					.1507	.1449		
					(.1091)	(.1238)		
$IPG(-4)$					−.2516	−.2025		
					(.1003)	(.1131)		
$\pi(-1)$							.4008	.3991
							(.1044)	(.1209)
$\pi(-2)$							.5520	.6162
							(.1112)	(.1268)
$\pi(-3)$							.1063	−.0179
							(.1119)	(.1275)
$\pi(-4)$							−.1837	−.1613
							(.1033)	(.1112)
R^2	.2952	.2023	.4563	.3766	.2148	.1496	.7305	.7555
SE	.0062	.0066	.0066	.0070	.0237	.0239	.0039	.0038
D-W	1.97	1.96	1.96	1.94	2.00	1.98	2.03	2.01

Note: Standard error of the coefficients in parentheses. Definitions of variables: $M1G$ = quarterly rate of growth of $M1$, $M2G$ = quarterly rate of growth of $M2$, IPG = quarterly rate of growth of industrial production, π = quarterly rate of growth of CPI, UN = unemployment rate, r = ninety-day bill rate, BOP = balance of payments on current account, G = quarterly rate of growth of real federal government expenditures, GDEBT = quarterly rate of growth of government debt, SURP = high employment surplus.

Table 5.A.2 **Multivariate Time-Series Models**

Model:	A2.1	A2.2	A2.3	A2.4	A2.5	A2.6	A2.7	A2.8
Dependent Variable:	*M1G*	*M1G*	*M2G*	*M2G*	*IPG*	*IPG*	π	π
Sample Period:	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976
Constant term	−.0004	.0015	.0026	.0044	−.0148	.0017	.0053	.0033
	(.0017)	(.0021)	(.0016)	(.0027)	(.0125)	(.0111)	(.0019)	(.0021)
$M1G(-1)$	.0906	−.1031			.6904			.1835
	(.1655)	(.1875)			(.3938)			(.0635)
$M1G(-2)$	.5233	.5336			.6883			.0376
	(.1680)	(.1906)			(.4089)			(.0649)
$M1G(-3)$	−.2765	−.3184			1.0207			−.0663
	(.1785)	(.1990)			(.4200)			(.0657)
$M1G(-4)$	−.1757	.0052			.2745			.1600
	(.1791)	(.1739)			(.4337)			(.0727)
$M2G(-1)$	.3590	.5998	.5050	.5211		.7132		
	(.1618)	(.1666)	(.1119)	(.1271)		(.3814)		
$M2G(-2)$	−.3972	.5612	.0985	.0955		.9460		
	(.1754)	(.1935)	(.1266)	(.1426)		(.4300)		
$M2G(-3)$	.3510	.4112	.1517	.1179		.7217		
	(.1757)	(.2077)	(.1167)	(.1325)		(.4122)		
$M2G(-4)$	.1396	.0240	−.0806	−.0838		−.2049		
	(.1645)	(.1843)	(.1036)	(.1180)		(.4182)		

Table 5.A.2 (continued)

Model:	A2.1	A2.2	A2.3	A2.4	A2.5	A2.6	A2.7	A2.8
Dependent Variable:	*M1G*	*M1G*	*M2G*	*M2G*	*IPG*	*IPG*	π	π
Sample Period:	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976	1954–1976	1959–1976
IPG(−1)					−.0947	−.2778		
					(.1622)	(.1129)		
IPG(−2)					−.6171	−.3626		
					(.1727)	(.1089)		
IPG(−3)					−.0902	−.0457		
					(.1709)	(.1089)		
IPG(−4)					−.2807	−.1869		
					(.1224)	(.1010)		
$\pi(-1)$					−.5539	−2.277	.2245	.1209
					(.6501)	(.793)	(.1079)	(.1273)
$\pi(-2)$					−.5186	−1.753	.5527	.7705
					(.6229)	(.701)	(.1053)	(.1171)
$\pi(-3)$					−.8774	.398	.2833	.3911
					(.6533)	(.846)	(.1090)	(.1339)
$\pi(-4)$					−.7464	1.252	−.0636	−.2328
					(.6533)	(.591)	(.1116)	(.1176)
UN(−1)					−.0904		−.0036	−.0052
					(.0096)		(.0009)	(.0014)
UN(−2)					−.0018		.0012	−.0016
					(.0120)		(.0016)	(.0021)
UN(−3)					.0173		.0018	.0034
					(.0114)		(.0015)	(.0021)
UN(−4)					−.0001		−.0004	−.0004
					(.0091)		(.0095)	(.0013)

$r(-1)$	−.2647		−.4966	−.5214		.9813		
	(.1067)		(.1000)	(.1221)		(.4501)		
$r(-2)$	.2086		.4853	.4827		.5253		
	(.1305)		(.1373)	(.1631)		(.4819)		
$r(-3)$	.1250		.0332	.0551		.3307		
	(.1396)		(.1472)	(.1669)		(.4793)		
$r(-4)$	−.0277		.0519	.0333		−1.830		
	(.1121)		(.1132)	(.1279)		(.437)		
$BOP(-1)$						.0058		.0001
						(.0035)		(.0006)
$BOP(-2)$						.0050		.0003
						(.0047)		(.0008)
$BOP(-3)$						.0016		−.0005
						(.0046)		(.0008)
$BOP(-4)$						−.0094		−.0010
						(.0037)		(.0006)
R^2	.5427	.4024	.6232	.5472	.5159	.7108	.7949	.8789
SE	.0053	.0059	.0056	.0061	.0200	.0160	.0035	.0029
D-W	1.97	2.01	1.96	1.95	2.03	2.00	1.96	2.21

Note: See table 5.A.1.

Table 5.A.3 ***F* Statistics for Significant Explanatory Power in Forecasting Equations of Four Lags of Each Variable**

	Model							
Variable	A2.1	A2.2	A2.3	A2.4	A2.5	A2.6	A2.7	A2.8
M1G	3.10*	2.44	2.32	2.17	6.53**	.09	1.66	4.14
M2G	3.98**	5.28**	15.37**	10.42**	.05	11.09**	1.12	1.13
IPG	.98	1.26	1.70	1.65	3.71**	4.70**	1.08	1.55
π	2.04	1.81	1.19	.40	7.61**	5.05**	79.32**	48.33**
UN	.77	1.42	.51	.25	3.00*	2.32	6.51**	5.88**
r	2.85*	2.46	9.19**	5.93**	1.95	5.62**	.69	.94
BOP	1.41	.79	1.04	.70	1.83	3.54*	2.32	2.87*
G	1.54	.47	.10	.22	.89	.46	2.25	1.59
GDEBT	.63	.43	.90	.44	.36	1.84	.57	1.33
SURP	.57	.62	1.09	1.28	.54	.70	1.57	2.01

Note: The *F* statistics test the null hypothesis that the coefficients on the four lags of each of these variables equals zero. The *F* statistics are distributed asymptotically as $F(4, x)$, where x runs from 47 to 83. The critical values of *F* at the 5 percent level are 2.5–2.6 and at the 10 percent level are 3.6–3.7.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

Table 5.A.4 **Chow Tests for the Models of Tables 5.A.1 and 5.A.2**

Model	F Statistic	Marginal Significance Level
A1.1	$F(5,82) = 3.50^{**}$	.0064
A1.2	$F(5,62) = 2.73^{*}$	.0271
A1.3	$F(5,82) = 2.49$	.0377
A1.4	$F(5,62) = 1.24$	.3014
A1.5	$F(5,82) = 2.81^{*}$	.0216
A1.6	$F(5,62) = 3.80^{**}$	.0046
A1.7	$F(5,82) = 1.59$	.1721
A1.8	$F(5,62) = 1.83$	.1200
A2.1	$F(13,66) = 1.81$	.0597
A2.2	$F(9,54) = 2.94^{**}$	.0065
A2.3	$F(9,74) = 1.60$	.1310
A2.4	$F(9,54) = 1.80$	.0895
A2.5	$F(17,58) = 1.39$	.1753
A2.6	$F(21,30) = 1.76$	.0765
A2.8	$F(17,38) = 1.24$	.2824

Note: Tests that the coefficients are equal in the two halves of the sample period.
*Significant at the 5 percent level.
**Significant at the 1 percent level.

Appendix 5.2: Additional Experiments Using the Two-Step Procedure

Because the two-step procedure used by Barro (1977) yields consistent parameter estimates and is far easier to execute than the joint nonlinear procedure used in the text, it is used in tables 5.A.5 and 5.A.6 to provide additional estimates of the long bond and short-rate models.

The two-step procedure here does not correct for heteroscedasticity within each long bond and short-rate equation even though Goldfeld-Quandt (1965) tests frequently reveal that it exists. This simplifies estimation and does not affect the consistency of the parameter estimates. However, this two-step procedure may yield incorrect standard errors and test statistics. Thus although tables 5.A.5 and 5.A.6 provide useful information, some caution about statistical inference is warranted.

The first four models of panels A and B in both tables reestimate the models in the text by the two-step procedure. As we might expect, the parameter estimates are similar to those génerated by the nonlinear procedures of the text and yield similar conclusions. This gives us some confidence that the two-step procedure can be used to gain further information on the robustness of this chapter's empirical results. Using the two-step procedure, long bond and short-rate models were also estimated with alternative specifications of the forecasting equations.

Table 5.A.5 **Estimates of the Long Bond Model: Using Seasonally Adjusted Data and the Two-Step Procedure (Sample Period 1954–1976)**

Model	Coefficients of							
	$M1G_t - M1G_t^e$	$M2G_t - M2G_t^e$	$MBG_t - MBG_t^e$	$URG_t - URG_t^e$	$UBG_t - UBG_t^e$	$IPG_t - IPG_t^e$	$\pi_t - \pi_t^e$	Constant Term
	Panel A: Using Residuals from Univariate Forecasting Models							
A5.1	.3507							−.0008
	(.5753)							(.0034)
A5.2	.3658					−.4447	−2.136	−.0008
	(.5460)					(.1412)	(.8580)	(.0033)
A5.3		1.113						−.0008
		(.530)						(.0034)
A5.4		.8762				−.4265	−1.987	−.0008
		(.5007)				(.1385)	(.8455)	(.0033)
A5.5			−.3502					−.0008
			(.5388)					(.0034)
A5.6			.1974			−.4409	−2.295	−.0008
			(.5195)			(.1418)	(.8793)	(.0033)
A5.7				.1955				−.0008
				(.1936)				(.0034)
A5.8				.2214		−.3952	−2.487	−.0008
				(.1909)		(.1437)	(.8788)	(.0033)
A5.9					.3703			−.0008
					(.4937)			(.0034)
A5.10					.4196	−.4064	−2.410	−.0008
					(.4879)	(.1436)	(.8796)	(.0033)

Panel B: Using Residuals from Multivariate Forecasting Models

A5.11	.2788							−.0008
	(.7149)							(.0034)
A5.12	.6262					−.5626	−1.997	−.0008
	(.6958)					(.1863)	(.994)	(.0033)
A5.13		1.105						−.0008
		(.642)						(.0034)
A5.14		1.197				−.5547	−1.973	−.0008
		(.608)				(.1795)	(.977)	(.0033)
A5.15			.3019					−.0008
			(.6563)					(.0034)
A5.16			.1308			−.5339	−2.077	−.0008
			(.6229)			(.1854)	(1.008)	(.0033)
A5.17				.1587				−.0008
				(.2204)				(.0034)
A5.18				.1558		−.5063	−2.195	−.0008
				(.2164)		(.1855)	(1.016)	(.0033)
A5.19					.3936			−.0008
					(.5875)			(.0034)
A5.20					.6392	−.5219	−2.311	−.0008
					(.5657)	(.1818)	(1.018)	(.0033)

Note: See table 5.A.1.

Table 5.A.6 **Estimates of the Short-Rate Model: Using Seasonally Adjusted Data and the Two-Step Procedure (Sample Period 1959–1976)**

Model	$M1G_t - M1G_t^e$	$M2G_t - M2G_t^e$	$MBG_t - MBG_t^e$	$URG_t - URG_t^e$	$UBG_t - UBG_t^e$	$IPG_t - IPG_t^e$	$\pi_t - \pi_t^e$	Constant Term	σ
	Coefficients of								
	Panel A: Using Residuals from Univariate Forecasting Models								
A6.1	.2777							.0002	−1.164
	(.1223)							(.0017)	(.287)
A6.2	.2930					.0538	.6534	.0001	−1.102
	(.1131)					(.0315)	(.1968)	(.0016)	(.266)
A6.3		.0766						.0000	−1.123
		(.1197)						(.0018)	(.298)
A6.4		.1632				.0615	.6723	.0002	−1.079
		(.1133)				(.0324)	(.2081)	(.0016)	(.276)
A6.5			.2291					−.0003	−1.078
			(.1287)					(.0017)	(.290)
A6.6			.1390			.0601	.5094	−.0008	−1.003
			(.1252)			(.0327)	(.2096)	(.0016)	(.277)
A6.7				.0095				−.0002	−1.093
				(.0450)				(.0017)	(.296)
A6.8				.0111		.0646	.5527	−.0008	−1.005
				(.0444)		(.0342)	(.2102)	(.0016)	(.280)
A6.9					.0752			−.0002	−1.097
					(.1157)			(.0016)	(.295)
A6.10					.0934	.0703	.5310	−.0008	−1.003
					(.1139)	(.0341)	(.2091)	(.0016)	(.278)

Panel B: Using Residuals from Multivariate Forecasting Models

A6.11	.1990							.0001	−1.130
	(.1442)							(.0015)	(.293)
A6.12	.2060					.0323	.3650	.0001	−1.139
	(.1451)					(.0609)	(.3147)	(.0018)	(.296)
A6.13		.0894						.0001	−1.092
		(.1397)						(.0017)	(.295)
A6.14		.1143				.0340	.3781	−.0001	−1.101
		(.1429)				(.0619)	(.3272)	(.0018)	(.298)
A6.15			.2054					−.0003	−1.056
			(.1617)					(.0017)	(.294)
A6.16			.1616			.0111	.6450	.0003	−1.126
			(.1569)			(.0427)	(.2398)	(.0017)	(.289)
A6.17				.0168				−.0001	−1.091
				(.0494)				(.0017)	(.296)
A6.18				−.0006		.0131	.6694	.0002	−1.158
				(.0549)		(.0435)	(.2443)	(.0017)	(.290)
A6.19					.1493			−.0004	−1.046
					(.1357)			(.0017)	(.296)
A6.20					.0901	.0161	.6377	−.0000	−1.121
					(.1325)	(.0435)	(.2443)	(.0017)	(.293)

Note: See table 5.A.1.

Models were estimated with residuals from the eighth-order autoregressive forecasting equations, as well as from multivariate models which included the four lagged values of a variable even if it was significant only at the 10 percent level (rather than the 5 percent level as in the text). The results were quite close to those reported here and the results again appear robust to changes in the specification of the forecasting equation.

Because the Federal Reserve might have changed its reaction function in the 1970s by paying more attention to monetary aggregates than it did previously, it is possible that the results here might substantially change if the 1970s are excluded from the sample period. Two-step estimates of the long bond and short-rate models over the sample period ending in the 1969:4 quarter fail to support this conjecture. The unanticipated IPG and π coefficients remain similar to those in tables 5.A.5 and 5.A.6: for the long bond model, the IPG coefficients range from $-.36$ to $-.46$ and the π coefficients from -1.69 to -2.17; and for the short-rate model, the IPG coefficients range from $-.04$ to $.03$ and the π coefficients from .37 to .57. Similar conclusions about the relationship of money growth and interest rates result also from estimates using the shorter sample periods. For the long bond model, the unanticipated $M1G$ coefficients are now negative, ranging from $-.26$ to $-.54$ and the $M2G$ coefficients range from .24 to .79. For the short-rate model, the money growth coefficients remain positive, with the $M1G$ coefficients ranging from .11 to .20 and the $M2G$ coefficients from .03 to .12.

The most obvious choice for the monetary aggregate that is exogenously determined by the Federal Reserve are not $M1$ and $M2$. As becomes clear from such debates as those between Anderson and Jordan (1969) and De Leeuw and Kalchenbrenner (1969), other aggregates may be a more sensible control variable for the Fed. If these aggregates are more likely than $M1$ or $M2$ to be exogenous, their use in the models here should give a clearer picture of the effect of monetary policy on interest rates. For this reason, tables 5.A.5 and 5.A.6 also contain two-step estimates of the models using the following additional variables:

MBG = growth rate of the monetary base (quarterly rate),
URG = growth rate of unborrowed reserves (quarterly rate),
UBG = growth rate of the unborrowed base (quarterly rate).

These variables are constructed analogously to $M1G$ and $M2G$ from the same data source, and the specifications for the forecasting equations were obtained with the same procedures used for $M1G$ and $M2G$.

In some applications the monetary base has been chosen as the Fed's control variable (e.g., see Anderson and Jordan 1968), while in monetary sectors of the large structural macroeconometric models such as the MPS (see Modigliani 1974) unborrowed reserves are often the exogenous control variable. On the other hand, the unborrowed base is the monetary aggregate corresponding most closely with open market operations. All three of the monetary aggregates are thus worthy candidates to be included in the long bond and short-rate models.

The results from using alternative monetary aggregates do not alter the conclusions or the relationship of monetary policy and interest rates. In the long bond models, the coefficients for the alternative aggregates are somewhat less positive than those for $M1$ or $M2$. They provide even less support for the view that an increase in monetary aggregates is associated with a fall in long interest rates. The coefficients in the short-rate models are almost always positive, and this is consistent with the results for $M1$ and $M2$, that a surprise increase in the monetary aggregate is associated with a rise in short rates.