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

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ABSTRACT

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, a hotly debated issue in the literature. 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. The empirical results uniformly *do not support* the proposition that increases in money growth are correlated with declines in short rates.

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 macromodels, has an increase in money growth leading, at least in the short and medium runs, to a decline in interest rates.[2] An alternative view, associated with Milton Friedman [6, 7], indicates that interest rates might rise in response to higher money growth because there is a rise in inflationary expectations and hence a rise in interest rates through a Fisher [5] 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 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]

The theory of efficient capital markets and rational expectations suggests an alternative approach for analyzing the relationship of money stock increases and

* University of Chicago. A more detailed version of this paper which contains additional results, Mishkin [12], is available as an NBER working paper. I thank James Pesando, Ed Burmeister, Ron Michener, Eugene Fama, and the participants in the Finance Workshop at the University of Chicago and the Money Workshops at the University of Chicago and the University of Virginia for their helpful comments. The National Science Foundation has provided research support. The usual disclaimer applies.

[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 [15].

[3] See Fama [2] and Mishkin [9].

[4] A more extensive discussion of the previous empirical work on this topic and references to this work can be found in Mishkin [11].

interest rate movements. A previous paper (Mishkin [11]) 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, allowing 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, which is estimated in the subsequent section. The paper then concludes with an interpretation of these results.

I. 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. More precisely, it implies that the market uses available information correctly in assessing the probability distribution of all future interest rates and hence:

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

where

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

ϕ_{t-1} = information available at time $t-1$,

$E(\cdots \mid \phi_{t-1})$ = the expectation conditional on ϕ_{t-1},

$E_m(\cdots \mid \phi_{t-1})$ = the market's expectations (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., $r_t^e = E_m(r_t \mid \phi_{t-1})$) then (1) implies

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

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

$$r_t - r_t^e = (X_t - X_t^e)\beta + \epsilon_t \tag{3}$$

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

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

β = a coefficient or vector of coefficients,

ϵ_t = serially uncorrelated error process (because $E(\epsilon_t \mid \phi_{t-1}) = 0$)

The rational expectations model (3) stresses that an unanticipated change[5] in the short rate will occur only when unanticipated information hits the market.

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 [3], 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.:

$$_{t-1}F_t = r_t^e + \delta_t \tag{4}$$

and

$$\delta_t = a_o + a_1\sigma_t \tag{5}$$

where

$_{t-1}F_t$ = 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:

$$r_t - {}_{t-1}F_t = -a_o - a_1\sigma_t + (X_t - X_t^e)\beta + \epsilon_t \tag{6}$$

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.[7] Substituting money growth for X_t in Equation (6) leads to[8]:

$$r_t - {}_{t-1}F_t = -a_o - a_1\sigma_t + \beta_m(MG_t - MG_t^e) + \epsilon_t \tag{7}$$

[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.

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

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

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] As has been found in foreign exchange markets (see, for example, Mussa [16]) quarterly changes in the spot rate, in this case the short rate, are primarily attributable to unanticipated movements in the spot rate. Also see Fama [3]. Using the model of the liquidity premium estimated in (9), 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.

[8] Other monetary aggregates are also potential candidates for inclusion in X_t. Of particular interest are such aggregates as unborrowed reserves or the unborrowed base which are under more direct control of the Federal Reserve and are thus possibly more exogenous than the M1 and M2 aggregates.

where

$$MG_t = \text{the money growth rate at time } t$$

The liquidity preference approach to the demand for money indicates that 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. 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:

$$r_t - {}_{t-1}F_t = -a_o - a_1\sigma_t + \beta_m(MG_t - MG_t^e) + \beta_y(YG_t - YG_t^e) + \beta_\pi(\pi_t - \pi_t^e) + \epsilon_t \quad (8)$$

where

$$YG_t = \text{growth rate of real income,}$$

$$\pi_t = \text{inflation rate,}$$

$$\beta_m, \beta_y, \beta_\pi = \text{coefficients}$$

This equation is really a rational expectations analog to the typical money demand relationship found in the literature.

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" macroeconometric models), the implication is that the coefficient on unanticipated money growth should be significantly negative in Equations (7) and (8): i.e., $\beta_m < 0$. The liquidity preference and Fisher [5] views lead us also to expect both positive income and inflation coefficients: i.e., $\beta_y > 0$ and $\beta_\pi > 0$.

Note that, as is pointed out in Mishkin [11], these coefficients are not invariant to changes in the time-series processes of money growth, income growth, and inflation. However, if, as is found in the sample period used here, these time-series processes display persistence—i.e., an unanticipated rise leads to a higher value in the coming period—then the expected signs are as above. Mishkin [11] also discusses another important caveat. The rational expectations model does not guarantee that Equations (7) and (8) are reduced forms where the right-hand-side unanticipated variables are uncorrelated with the error term, ϵ_t, and therefore the estimates of β are consistent.[9] Because the analysis in this paper cannot demonstrate that Equations (7) and (8) are reduced forms, the β estimates should be viewed only as providing information on the correlations of such variables as unanticipated money growth with the unanticipated change in short rates. Interpretation of these correlations is deferred to the concluding remarks.

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

As is discussed later, exogeneity of the monetary aggregate is desirable because interpretation of results is then more clear-cut. Therefore, in future research I plan to estimate the models with these other aggregates.

[9] This issue of the consistency of the β estimates is discussed more extensively in Abel and Mishkin [1].

II. 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. Since misleading results can be obtained from efficient markets models using averaged data,[10] the interest rate data 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 rather than the GNP deflator has been used to calculate the inflation variable.[11] A more detailed description of the data is available in a longer version of this paper, Mishkin [12].

The Results

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 [3] 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 $t - 1$ poses 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., $r_t - {}_{t-1}F_t$, 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

$$r_t - {}_{t-1}F_t = -.0001 \quad 1.0961\,\sigma_t + \epsilon_t$$

$$-$$

$$\qquad (.0017) \quad (.2937)$$

$$R^2 = .1659 \qquad \text{Standard Error} = .0068 \qquad \text{Durbin-Watson} = 1.90 \qquad (9)$$

where

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

As in Fama [3], 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

[10] See Working [19], for example.

[11] In previous work (Mishkin [11]), 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.

used in the empirical tests that follow. However, the specification of the risk premium model is not a critical issue to the outcomes: use of a Fama measure of σ_t or the exclusion altogether of σ_t from the model does not alter the results appreciably.

The (7) and (8) models are estimated jointly with forecasting equations for money growth, industrial production growth and inflation, imposing cross-equation constraints implied by rationality of expectations.[12] The estimation procedure is discussed in Mishkin [11] and [13]. In order to verify the robustness of the results, two methods, rather than one, are used for specification of the forecasting equations, and they are discussed in Mishkin [11]. The simplest specifies these equations as univariate autoregressive time-series models, while the other method involves specifying multivariate models using Granger's [8] concept of predictive content.[13] 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 with these data appear in Table I. Panel A of this table contains estimates where the forecasting equations are univariate time-series models, while for Panel B the forecasting equations are multivariate models.

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.[14] Not only are both of these coefficients in Models 1.1 and 1.2 positive rather than negative, but they are also significantly different from zero at the 1% level. Nor are the coefficients numerically small. They indicate that a 1% surprise increase in M1 is associated with a 25 to 30 basis point unanticipated increase in the bill rate.[15] The Panel B estimates of the M1G coefficients indicate

[12] To be more precise, the nonlinear constraints are generated by two hypotheses: (1) rational expectations, and (2) the model of market equilibrium in Equations (4) and (5). Likelihood ratio tests reported in the longer version of this paper, Mishkin [12], indicate that in over half the models these constraints are rejected at the 5% level. As is argued in Mishkin [12], it is very plausible that these rejections are due to the inadequacy of the model of market equilibrium rather than the nonrationality of expectations. Then it is easily shown that the rational expectations models estimated here are still a valid framework for analyzing the relationship of money growth and short interest rates. Incorrect specification of the model of the risk premium will not lead to inconsistent estimates of the β coefficients. Since the derivation of a better model of the risk premium is not necessary for achieving reliable estimates of the β's if expectations are rational, this tricky research issue, which is beyond the scope of this paper, is left as a subject for further research.

The results in the longer version of this paper do provide some information on the possible direction of this research. Despite Fama's [3] finding that his measure of the risk premium is successful in passing the market efficiency test that forward rates adjusted for this measure outperform a random walk scheme in forecasting interest rates, a more powerful test, such as here, demonstrates that this measure is not adequate. This should not be surprising because there is no obvious utility function that would lead us to this measure. A theoretically more justifiable risk measure would exploit the covariance of treasury bill yields with returns on alternative assets. In addition, these results indicate that the risk premium is possibly related to past money growth and the level of interest rates, while they give no indication that it is related to past inflation, industrial production growth, unemployment or the balance of payments.

[13] The specifications for these equations can be found in Tables 1 and 2 of Mishkin [12].

[14] Urich and Wachtel [18] obtain a similar result using weekly data. Reduction in the unit of observation in the analysis here is therefore likely to leave the findings intact.

[15] 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%.

Table I

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 Forecasting Equations						
1.1	.3029**				.0003	−1.1255**
	(.0652)				(.0014)	(.2530)
1.2	.2458**		.0274	.4687**	.0001	−1.1267**
	(.0671)		(.0171)	(.1716)	(.0014)	(.2464)
1.3		.1926*			.0003	−1.1468**
		(.0644)			(.0015)	(.2765)
1.4		.1967**	.0440*	.5459**	.0001	−1.1260**
		(.0624)	(.0176)	(.1746)	(.0014)	(.2526)
Panel B. Using Multivariate Forecasting Equations						
1.5	.3431**				.0007	−1.2403**
	(.0831)				(.0015)	(.2639)
1.6	.2484**		.0386	.5079	.0004	−1.2037**
	(.0956)		(.0376)	(.2623)	(.0016)	(.2861)
1.7		.3285**			−.0007	−.9891**
		(.0918)			(.0015)	(.2841)
1.8		.2011*	.0400	.5788*	−.0003	−1.0791**
		(.0986)	(.0374)	(.2674)	(.0016)	(.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 parentheses.
$M1G-M1G^e$ = unanticipated M1 growth
$M2G-M2G^e$ = unanticipated M2 growth
$IPG-IPG^e$ = unanticipated industrial production growth
$\pi-\pi^e$ = unanticipated inflation
σ = measure of interest rate uncertainty

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 and significantly different from zero at the 1% level. The coefficients on unanticipated M2 growth tell a similar story. They also display a significant positive correlation of unanticipated money growth and unanticipated bill rate movements.

How different are the results found here from those that might be inferred from "Keynesian" structural macroeconometric models? The response of one such model, the MPS (MIT-Penn-SSRC) Quarterly Econometric Model [14], to a 1% surprise increase in M1 growth has been analyzed with a simulation technique discussed in Mishkin [10].[16] The MPS model indicated that this 1% M1 surprise led to an immediate *decline* of 88 basis points in the bill rate. This strongly contrasts with the finding here that even the least positive M1 coefficient is more than ten 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.[17] 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 [4] that forecasts from univariate time-series models may be "economically rational" expectations. The results on the unanticipated inflation and industrial production coefficients in Panel A do conform to our priors. In both the M1 and M2 rational expectations models, these coefficients are positive and the inflation coefficients are significantly different from zero at the 1% level while only one of the industrial production coefficients is significantly different from zero at the 5% 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 rational expectations model does not specify whether the $X - X^e$ variables should be described by seasonally adjusted or 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 rational expectations models have also been estimated using seasonally adjusted data rather than unadjusted data as in Table I. For the sake of brevity, these results are not reported here, but they can be found in Mishkin [12]. The coefficient estimates are very similar, but their standard errors are somewhat larger. We continue to find that all the money growth coefficients are positive with three out of eight statistically significant. The seasonally adjusted data also do not support the view that an unanticipated increase in money growth is associated with an unanticipated fall in short rates. However, they are not quite as strong as the unadjusted data in supporting the opposite view.

[16] More details on this simulation experiment are given in Mishkin [11].

[17] Additional results in the Appendix of Mishkin [12] also support this view. As Feige and Pearce [4] 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 Panels A and B gives some credibility to this viewpoint and this issue is discussed more extensively in Mishkin [13].

III. Conclusion

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 longer version, Mishkin [12]. 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 [11], which conducts a parallel analysis of long-term interest rate behavior.

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 and a resulting lack of correlation of unanticipated money growth with the error term in the rational expectations model, 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. However, if unanticipated money growth is correlated with the error term, then the β_m coefficient estimates are inconsistent and can result in 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 [8] "causality" running from interest rates to money growth does not imply that the estimates of β_m will be inconsistent. Tests of the Sims [17] 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 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.

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