

A TEST OF FISCHER'S THEORY OF MONETARY MISPERCEPTIONS AND THE BUSINESS CYCLE IN THE PRESENCE OF LONG-TERM CONTRACTS

Charles W. Bischoff

State University of New York at Binghamton

and

Steven C. Hine

State University of New York at Binghamton

INTRODUCTION

Many responses have been presented over the past fifteen years to the radical "policy ineffectiveness proposition" of the new classical school. The proposition implies that deviations of real economic activity from trend depend on deviations of actual from anticipated money over short periods of time but not on the anticipated part of money. One such early and influential response was that of Stanley Fischer [1977, 1980] wherein he demonstrates that the existence of multi-period wage contracts is sufficient to re-introduce an effective role for stabilization policy even if expectations are formed rationally¹. Despite the importance of his work, however, there exists no direct attempt to test the key implications of Fischer's model even after more than a decade since its publication.

Perhaps the reason for the lack of such an attempt is the scarcity of necessary appropriate data; in Fischer's model, deviations of output from its natural rate depend on deviations of actual money from expected values formed several periods earlier. Indeed, fluctuations in output depend on expectational errors going back to the point in time at which the longest existing contract was drawn up. Unfortunately, such multi-period-ahead expectations pertaining to monetary aggregates are not readily available in the United States. There are a few such series, however, which have been formulated as necessary inputs into the operation of some large-scale structural forecasting models which date back to about 1970. One innovative aspect of this paper is the use of one of these series to estimate and test the explanatory power of a version of Fischer's multi-period contracting model.

It is not enough, of course, to merely confront a plausible theory with the data by estimating the parameters of a model and considering its fit or significance level and the sign of various coefficients. A more persuasive approach is to subject the new model to a statistical test of its explanatory power relative to or in the presence of a leading alternative². For the purpose of providing just such an alternative, we employ a widely-known version of Lucas's model based on one-period-ahead monetary misperceptions, as presented by Barro [1977, 1978] and Barro and Rush [1980]³.

In the following section, the two competing specifications for output and unemployment determination are presented in more detail. A non-nested model testing procedure appropriate for the particular models considered here is also put forth. The next section contains empirical results of the estimation and testing procedure. Monte Carlo evidence

regarding the small sample properties of the test statistic used is also presented. Lastly, there is a section summarizing the results.

MODEL SPECIFICATIONS AND EVALUATION PROCEDURE

This study first represents the Barro-type specifications for quarterly rates of output and unemployment as equations which include a distributed lag of eight one-quarter-ahead money growth forecast errors ($m_{t,i} - E_{t-1} m_{t,i}$ for $i=0, \dots, 7$) determining the dependent variable⁴. Then, following Fischer's suggestion, we represent the competing specifications as equations having output and unemployment as functions of the one- through eight-quarter-ahead money growth forecast errors ($m_t - E_{t-j} m_t$ for $j=0, \dots, 7$)⁵. The distinction between the Barro and Fischer specifications is critical for, as Fischer states,

...it is very hard to argue that the Fed cannot use a monetary rule that reacts within a period of two years to new information. If the two-year expectation is somehow locked in (for example, in labor contracts), then the Fed has ample time to act to affect the behavior of output. That does not mean it *should* act, but rather that it can systematically affect output. [1980, 220]

Thus it is anticipated that by reconsidering the performance of these two model specifications in light of newly available empirical techniques, we might further our understanding of the roles that both nominal contracting and informational confusion play in the generation of the business cycle.

Fischer did attempt to provide support for his contracting-based model of output by adding a measure of unanticipated money growth over two years to Barro's yearly output equation. He observes that "adding the variable ... to the Barro equation reduces the sum of squared residuals" by about 12 percent, while "if the current value of the DMR variable is then deleted from the regression, the sum of squared residuals rises only slightly" by about 1 percent. While this evidence is suggestive of the importance of nominal rigidities, formal F-tests still lead Fischer to "conclude that the data cannot tell us whether only one-year ahead or only two-year ahead errors in predicting money or both contribute to explaining the behavior of output." [1980, 235] This is, as Fischer is aware, due to the inability to identify any significant independent effect that the various expectational errors have on output. Thus, alternative techniques are needed if a conclusive evaluation of the relative significance of such competing explanations is to be conducted.

Davidson and MacKinnon [1981], Bernanke, Bohn and Reiss [1988] and many others have developed and refined procedures first proposed by Cox [1961, 1962] and implemented by Pesaran [1974], to test a specified model when there exists a non-nested alternative. This paper employs a slight generalization of Bernanke, Bohn and Reiss's [1988] extension of Davidson and MacKinnon's [1981] P-test to the situation where errors are serially correlated (yielding their P2-test). Our modification is necessitated by the presence of second-order serial correlation in both Barro and Rush's [1980] results and our estimates of identical equations on different data.

In order to define terms and describe our empirical procedure, suppose we take as the null hypothesis Barro's model of output, written as

$$(1) \quad y_t = f(X_t, \alpha) + \mu_{0t}$$

where y_t denotes the log of output in quarter t and where

$$(2a) \quad f(X_t, \alpha) = \alpha_0 + \alpha_1 t + \alpha_2 G_t + \sum_{i=0}^7 \alpha_{i+3} (m_{t,i} - E_{t-1} m_{t,i}) \quad \text{and}$$

$$(2b) \quad \mu_{0t} = p_{01} \mu_{0t-1} + p_{02} \mu_{0t-2} + \epsilon_{0t}$$

with G_t representing the log of real federal purchases of goods and services and t a time trend⁶. The error term, ϵ_{0t} , like ϵ_{1t} and e_t below, is assumed to be normally and independently distributed with a zero mean and constant variance. As the competing alternative, we express Fischer's model as

$$(3) \quad y_t = g(Z_t, \beta) + \mu_{1t}$$

where

$$(4a) \quad g(Z_t, \beta) = \beta_0 + \beta_1 t + \beta_2 G_t + \sum_{j=0}^7 \beta_{j+3} (m_t - E_{t-j} m_t) \quad \text{and}$$

$$(4b) \quad \mu_{1t} = p_{11} \mu_{1t-1} + p_{12} \mu_{1t-2} + \epsilon_{1t}$$

Letting carets (^) denote non-linear least-square estimates⁷, the Davidson-MacKinnon P-test is conducted by estimating the auxiliary equation

$$(5) \quad y_t - f(X_t, \hat{\alpha}) = \delta [g(Z_t, \hat{\beta}) - f(X_t, \hat{\alpha})] + \hat{F}_t' b + e_t$$

where \hat{F}_t denotes the derivatives of the model under the null (Barro's) hypothesis with respect to the α_i 's and p_{oi} 's, evaluated at the least-square estimates. More explicitly, if we express the Barro output model given above as

$$(6) \quad y_t = \alpha_0 + \alpha_1 t + \alpha_2 G_t + \sum_{i=0}^7 \alpha_{i+3} (m_{t,i} - E_{t-1} m_{t,i}) + p_{01} [y_{t-1} - \alpha_0 - \alpha_1(t-1) - \alpha_2 G_{t-1} - \sum_{i=0}^7 \alpha_{i+3} (m_{t-1,i} - E_{t-2} m_{t-1,i})] + p_{02} [y_{t-2} - \alpha_0 - \alpha_1(t-2) - \alpha_2 G_{t-2} - \sum_{i=0}^7 \alpha_{i+3} (m_{t-2,i} - E_{t-3} m_{t-2,i})] + \epsilon_{0t}$$

we can see that the components of the vector \hat{F}_t include:

$$(7a) \quad dy_t/d\alpha_0 = 1 - \hat{p}_{01} - \hat{p}_{02}$$

$$(7b) \quad dy_t/d\alpha_1 = t - \hat{p}_{01}(t-1) - \hat{p}_{02}(t-2)$$

$$(7c) \quad dy_t/d\alpha_2 = G_t - \hat{p}_{01} G_{t-1} - \hat{p}_{02} G_{t-2}$$

$$(7d) \quad dy_t/d\alpha_{i+3} = m_{t,i} - E_{t-1} m_{t,i} - \hat{p}_{01} (m_{t-1,i} - E_{t-2} m_{t-1,i}) - \hat{p}_{02} (m_{t-2,i} - E_{t-3} m_{t-2,i}), \quad i = 0, \dots, 7$$

$$(7e) \quad dy_i/dp_{0j} = y_{t,j} - \hat{\alpha}_0 - \hat{\alpha}_1(t-j) - \hat{\alpha}_2 G_{t,j} - \sum_{i=0}^7 \hat{\alpha}_{t+3} (m_{t+i,j} - E_{t+i-1} m_{t+i,j}), \quad j = 1, 2.$$

The P-test procedure now calls for testing the validity of the null hypothesis by considering whether $\hat{\delta}$ is significantly different from zero. Davidson and MacKinnon [1981] prove that if equations (1) and (2) represent the *true* model, then the distribution of the ratio of the estimated coefficient $\hat{\delta}$ to its standard error converges to the standard normal distribution as the sample size increases indefinitely. If the t-ratio for $\hat{\delta}$ exceeds 1.96, we would then normally conclude that the null hypothesis embodied in equations (1) and (2) is rejected in favor of the alternative specified by equations (3) and (4).

As Davidson and MacKinnon [1981] point out, a test of the Fischer model can be carried out against the Barro alternative by simply reversing the role that each model plays in the above procedure. In the reversal, a significant estimate of δ is an indication that one should favor the informational confusion explanation of the business cycle over the multi-period contracts story. If $\hat{\delta}$ is not significant in either equation, the result is akin to collinearity; the data do not suffice to distinguish between the two models. If $\hat{\delta}$ is significant in both cases, the result may be a spurious one due to small sample biases in the testing procedure (see below) or it may be that the true model contains elements of each of the competing alternatives.

This test procedure is, as are most specification tests for non-nested models⁸, an asymptotic test with small sample properties which are generally unknown. However, a number of studies⁹ of the small-sample properties of these tests find, on the basis of Monte Carlo simulations, a prevalent bias towards too high a probability of rejecting a true null hypothesis; that is, the probability of type I error or the size of the test is high relative to the asymptotic or nominal size.

We conducted Monte Carlo simulations designed to generate evidence on the significance levels associated with those tests which, on the basis of an asymptotic critical value of 1.96, lead us to reject the null hypothesis. Paralleling the method used by Bernanke, et.al. [1988], we used the estimated parameters, including the autoregressive coefficients and the standard error, and the explanatory variables of the null hypothesis to generate 1000 normally-distributed independent sample realizations for the dependent variable in each case studied. These sample realizations were then used to conduct the P2-test as outlined above, thereby generating 1000 values for our test statistic. While the empirical distribution of these values should then approximate the small-sample distribution of the test statistic under a true null, our concern was only with the properties of this distribution at the upper tail¹⁰. As a summary of the small sample size of our test, the proportion of times the simulation yields a value for the t-statistic which is greater than that which we actually found is presented. This we interpret as being the estimated probability of attaining as high a t-statistic value as we got when the null hypothesis is true. If this probability is less than 0.05, then we conclude that the alternative specification under consideration does indeed reject the null hypothesis at the standard level of confidence.

Additional evidence bearing on the issue at hand was generated by performing the above procedure using unemployment as well as the measure of real economic activity.¹¹ Also, the tests were conducted on both output and unemployment models using data on the *natural rate* for each of these real aggregates¹² rather than implicitly assuming a linear time-trend for potential output and a constant natural rate of unemployment.

EMPIRICAL RESULTS

Data on projected money supply levels for the current period through seven quarters ahead were available for the period 1970:2 through 1979:4¹³. From this, annualized expected growth rates for the next k quarters, for $k = 0, \dots, 7$, were generated for each of the sample periods¹⁴, thus allowing us to estimate the models over the period 1972:3 to 1979:4 (seven lags on the expectational errors were used as well as the second-order autoregressive error process, resulting in nine periods being excluded from the beginning of the sample). Actual money growth rates for corresponding time spans were based on revised data¹⁵, and k -period ahead expectational errors (at annual rates) were then the difference between actual and expected money growth rates.

The Chase measures of expectational errors may be criticized as being unrepresentative of aggregate errors since they are realizations of a sample of size one. However, we feel the use of these errors offers certain advantages over alternative measures. The econometric difficulties that arise with generated errors such as those used by Barro and Rush are well-documented¹⁶, but above all, the use of actual forecasts allows for the possibility that shifts in the expectations formulation process occur during our sample period. Thus Lucas's [1976] critique of the econometric estimation of models involving rational expectations is less likely to be germane. Furthermore, an empirical comparison of the one-period-ahead errors used here to those contained in Barro and Rush [Table 2.3] favors the Chase measure in terms of the magnitude of these errors. For the period 1970:2 to 1978:1, the period for which we have overlapping measures on both errors, in terms of annualized percentage growth rates, both the root-mean-squared errors (1.9% vs. 2.1%) and mean absolute errors (1.5% vs. 1.6%) indicate that the Chase forecasts performed better than the equation used by Barro and Rush to generate their DMR series. In addition, in a study comparing the forecasts of Chase, Data Resources, Inc., and Wharton EFA, McNees [1979, 41] finds that in terms of the mean absolute errors in current dollar terms, Chase does as well as DRI and better than Wharton. Thus, for both methodological and empirical reasons, the results found here may be more credible by virtue of our use of these data on expectations than if alternative measures had been employed.

The other data required, y , U and G , are taken to be U.S. figures for the log of real GNP (in 1972 dollars), the total labor force unemployment rate, and the log of real federal government purchases of goods and services, in order to approximate Barro and Rush's [1980] specification as closely as possible. Results of the estimation of the two competing models are given in Table 1 (using the original specification of the equations) and Table 2 (using the Bureau of Economic Analysis's measures of the natural rates).

It should be noted that the results in the first column of Table 1 are very comparable to those in Barro and Rush's [1980] Table 2.1, Column (5), both in terms of the size of the coefficient estimates and the estimate of the standard error of the regression, despite the shorter sample period.¹⁷ In addition, the use of actual measures of the natural rate do not noticeably alter the estimation results for either of the models of output. Although some point estimates are affected in the unemployment models, the pattern on the money error terms and the overall fit of the models are not.

In all four versions of the Fischer model, the largest coefficients appear on expectational errors dating back six and seven quarters. This means that misperceptions of money growth rates 17 to 20 months ahead are the most important factors affecting real variables. This leaves plenty of time for announced policy actions (and anticipated

TABLE 1
Models Estimated on Original Data

Variable	Output Models		Unemployment Models	
	Barro Model	Fischer Model	Barro Model	Fischer model
c	6.97 (142.8)	7.00 (353.8)	-3.28 (-6.9)	-2.66(-10.24)
t	0.008 (4.53)	0.007 (9.45)		
G	0.069 (0.59)	0.081 (1.24)		
G/y			7.68 (1.29)	-0.93 (-0.29)
um0	0.07 (2.42)	-0.14 (-2.92)	0.09 (0.14)	0.63 (3.61)
um1	0.25 (1.22)	0.03 (0.46)	-0.56 (-0.53)	0.49 (1.96)
um2	0.26 (1.00)	0.08 (0.8)	-0.99 (-0.71)	0.05 (0.13)
um3	0.28 (0.88)	-0.29 (-1.75)	-0.84 (-0.49)	2.47 (3.85)
um4	0.17 (0.51)	0.08 (0.41)	-0.35 (-0.2)	-0.01 (-0.01)
um5	0.01 (0.04)	0.47 (1.91)	0.47 (0.32)	-3.94 (-4.02)
um6	0.13 (0.53)	1.13 (4.54)	0.47 (0.4)	-5.46 (-5.79)
um7	0.01 (0.08)	-0.07 (-0.27)	0.27 (0.44)	-1.38 (-1.42)
p ₁	1.37 (6.39)	1.56 (9.71)	1.62 (10.28)	1.78 (16.81)
p ₂	-0.55 (-2.41)	-0.76 (-4.64)	-0.73 (-4.54)	-0.9 (-8.62)
s.e.	0.010754	0.007097	0.049733	0.030066
DW	2.14	2.39	1.57	1.37

t - statistics in parentheses

columns corresponding to Barro models, $um_i = m_{t,i} - E_{t,i-1}m_{t,i}$, $i=0,\dots,7$

columns corresponding to Fischer models, $um_j = m_t - E_{t,j-1}m_t$, $j=0,\dots,7$

s.e. = standard error of the regression

DW = Durbin-Watson Statistic

money changes) to affect output and unemployment.

The coefficient on unexpected money growth based on expectations formed two months earlier is positive and statistically significant at conventional levels in both the unemployment equation versions of the Fischer model and is negative in both, and significant in one of, the output versions. This implies that over this time period unexpected increases in money growth reduce real economic activity. There is a substantial literature linking unexpected rises in money to very rapid rises in interest rates during both the 1970s and the 1980s [Urich and Wachtel, 1981; Cornell, 1983a, 1983b; and Nichols, Small, and Webster, 1983], although there is no agreement on the mechanism except that it involves expectations. A linkage such as this could explain our coefficients if increases in unexpected money raised real interest rates and thus decreased output via a *Keynesian* mechanism. This phenomenon is not a part of the Fischer model but neither is it necessarily in conflict with it. In any case, the impact elasticity of either output or unemployment with respect to money in the Fischer model is the sum of the eight coefficients on the unexpected money growth terms. These elasticities are 1.29 and 1.37 for the output equations and -7.15 and -9.63 for the unemployment equations. The likelihood ratio test statistics for the hypothesis that the sums are zero are 15.17,

TABLE 2
Models Estimated on BEA Data

Variables	Output Models		Unemployment Models	
	Barro Model	Fischer Model	Barro Model	Fischer Model
G	0.107 (1.11)	0.131 (2.15)		
G/y			0.04 (0.04)	-0.67 (-1.43)
um0	0.12 (1.12)	-0.07 (-1.34)	-0.07 (-0.14)	0.48 (2.25)
um1	0.26 (1.61)	0.08 (1.13)	-0.35 (-0.37)	0.75 (2.42)
um2	0.18 (0.9)	0.08 (0.77)	-0.58 (-0.47)	0.85 (1.86)
um3	0.23 (0.99)	-0.29 (-1.63)	-0.48 (-0.32)	1.37 (1.88)
um4	0.12 (0.5)	0.16 (0.74)	0.54 (0.36)	-1.18 (-1.3)
um5	-0.01 (-0.03)	0.56 (2.09)	1.22 (0.95)	-3.08 (-2.66)
um6	0.12 (0.61)	0.87 (3.36)	0.54 (0.54)	-6.43 (-5.59)
um7	0 (0)	-0.02 (-0.06)	0.17 (0.32)	-2.39 (-1.97)
P ₁	1.4 (6.96)	1.43 (8.25)	1.7 (11.9)	1.73 (14.16)
P ₂	-0.54 (-2.68)	-0.66 (-3.72)	-0.82 (-5.53)	-0.91 (-7.32)
s.e	0.010007	0.007437	0.050253	0.035965
DW	2.14	2.21	1.94	1.62

See notes from Table 1.

TABLE 3
P-Test Results

Model	Barro null vs. Fischer alternative	Fischer null vs. Barro alternative
Output ¹	0.986 (4.78)	0.212 (0.5)
Unemployment ¹	0.998 (5.73)	0.238 (0.54)
Output ²	0.927 (4.19)	0.394 (1.07)
Unemployment ²	0.914 (5.16)	0.606 (2.09)

1. Models estimated on original data.

2. Models estimated on Bureau of Economic Analysis data.

14.79, 22.78, and 23.70 respectively. Under the null hypothesis, these statistics would be asymptotically distributed as chi-square with one degree of freedom. The fact that the coefficients often change signs as the horizon changes is likely to be attributable to multicollinearity, as the multi-period-ahead expectational errors overlap each other.

In addition, we may notice that in all cases the multi-period contracting model has a

TABLE 4¹⁸
Monte Carlo Results

Model	Barro null vs. Fischer alternative	Fischer null vs. Barro alternative
Output ¹	0.025	NA
Unemployment ¹	0.006	NA
Output ²	0.047	NA
Unemployment ²	0.009	0.302

1. Models estimated on original data.

2. Models estimated on Bureau of Economic Analysis data.

lower standard error than does the comparable new classical specification, which is suggestive evidence in favor of these models, but which is also the sort of claim which needs to be supported by exacting statistical tests. This was the purpose of the next step of the procedure.

Each of the four sets of competing models was used to generate the transformed variables needed to conduct the P-test according to equations (7a) through (7e) above. This was done twice, first taking the Barro-type specification as the null hypothesis and Fischer's multi-period contracting model as the alternative, and then again with the roles reversed. The estimates of and corresponding t-statistics (in parentheses) for each of the eight tests are given in Table 3.

If these t-ratios actually could be treated under the null hypothesis as random draws from a standard normal distribution, they would constitute strong evidence for rejection of the Barro model (though the Fischer model would also be rejected in one case out of four). However, the Monte Carlo studies cited in note 9 above strongly suggest that with a small sample such as this one the test statistic may not be treated this way. Thus, we turn to our simulation analysis. Table 4 presents the relative frequency with which the Monte Carlo simulations rejected a true null hypothesis with t-ratios greater than those in Table 3. These values thus represent the estimated area in the upper tail of the distribution of our test statistic under a true null hypothesis. If the statistic in Table 4 is less than .05, the corresponding test statistic in Table 3 is large enough that we estimate it would occur by chance less than five percent of the time if the null hypothesis were true. The only case of an apparent false rejection of a true null is the one where the Barro alternative rejects the Fischer null. We estimate that in that particular case, there is a probability of 30.2 percent of observing a t-statistic larger than 2.28 when in fact the Fischer model is true. All other cases yield estimates of the probability of a false rejection of a true null below 5 percent, the conventional level of significance. Indeed this evidence suggests that in both of the unemployment equation cases, the rejection by Fischer's model of the Barro model achieves a 99 percent level of confidence.

On the basis of their ability to explain the residuals of the alternatives, these results strongly favor the model which explicitly accounts for the existence of multi-period contracts, over an important and well-known alternative. This conclusion is significant

as it represents not only the first attempt to confront the Fischer model with data, but also the first evidence that such a model explains the data quite well. As the theoretical justification for this empirical specification has been clearly articulated by Fischer and others, the results here also emphasize the importance of nominal rigidities as a mechanism by which monetary changes are transmitted to the real economy.

SUMMARY AND CONCLUSIONS

We have utilized multi-period *ex ante* expectations data on the old M1 money supply to estimate the parameters of a model, suggested by Fischer, in which money affects real variables only through multi-period errors in anticipations. We have tested this model against an alternative, first tested by Barro, in which money affects real variables only through current and lagged single-period errors in anticipations. Previous work on the Barro model has involved the use of a second equation to generate monetary anticipations, but we have used *ex ante* anticipations for both models.

The test we used is a version of Davidson and MacKinnon's P test, in which a non-linear equation is linearized and then the predictions from an alternative model are added as a single extra variable. In our case, the non-linearity in the model arises because of second order serial correlation of the errors. Because we suspect that the small sample properties of the test lead to excessive rejection of the null hypothesis, we carried out random sampling experiments. These were intended to tell us whether the test statistics that we obtained are large enough to be unusual if the null hypothesis were true. We found that it is possible to reject the Barro model at conventional levels of confidence. However, we also found that when the Fischer model is used as the null hypothesis, it can not be rejected at conventional levels. We conclude that this evidence indicates that the Fischer model is superior to the Barro model. That is, multi-period expectational errors which capture the effects of long-term contracts represent important explanatory variables in models purporting to explain the business cycle. This conclusion supports the notion that systematic monetary policy can have an affect on real variables, although it does not address the issue of whether it should be used.

We would like to thank Robert Basmann for comments on an earlier draft.

NOTES

1. Other authors have presented similar arguments based on various nominal rigidities, including Phelps and Taylor [1977] and Taylor [1979, 1980].
2. As Bernanke, Bohn, and Reiss [1988] state "...in addition to or in place of the discrimination criteria used before, specifications should be evaluated by how well they can explain misspecifications in alternative models." [1988, 294].
3. The approach that Barro initiated in this series of papers and the results presented therein have been attacked on both theoretical and empirical grounds. We do not find the evidence presented by some authors to empirically refute the Barro and Rush [1980] study to be conclusive. A discussion of some of the most well-known of these studies is contained in an appendix available from the authors. In any case, the Barro and Rush model is an important and attractive alternative, especially in light of the difference between its implications regarding the efficacy of stabilization policy and those of the Fischer model. *Keynesian* alternatives are usually presented in multi-equation form, although Pesaran [1982a, 1988]

and Rush and Waldo [1988] conduct comparisons of single-equation specifications of Keynesian and new classical models. Nonetheless, it is not clear that Keynesian specifications represent an alternative to the Fischer model in the sense that they necessarily conflict. *Real Business Cycle* models, on the other hand, are not generally presented in regression form but rather are calibrated and their ability to simulate actual data then assessed.

4. The notation $m_{t,i}$ represents the difference between the log of money in periods $t-i-1$ and $t-i$. The term $E_{t-i-1}m_{t,i}$ represents the expectation formed in period $t-i-1$ of $m_{t,i}$. Barro [1978] refers to Lucas [1975] and a draft version of Sargent [1978] to justify the use of lagged values of the difference between actual and anticipated money growth to explain current real variables.
5. Given Barro's rationalization mentioned in the above note for the appearance of the distributed lag of expectational errors, we could consider a generalization of the Fischer specification

which contains $\sum_{j=0}^m \sum_{i=0}^n \alpha_j \beta_i (m_{t,i} - E_{t-i-1}m_{t,i})$. This specification is recognized by Fischer himself

[1977, 203, fn. 20]. However, when we tested this model with n equal to 2, a log-likelihood ratio test of the joint significance of the two additional coefficients failed to reject the hypothesis that they were zero. Thus the simpler specification is retained.

6. Equations (1) and (2) correspond to the specification reported by Barro and Rush [1980, 36] in Table 2.1, Column (5). The specification in Table 2.1, Column (4), could not be used because the variable MIL_t is zero throughout our sample period. The equation with unemployment as the dependent variable, mentioned below, corresponds to Barro and Rush's Table 2.1, Column (9).
7. Regressions with serially correlated errors were estimated using the Cochrane-Orcutt procedure with Micro TSP Version 6.5.
8. See McAleer [1987] for a thorough discussion of the various other non-nested tests and their properties.
9. These include Pesaran [1974, 1982b], Davidson and MacKinnon [1982], Godfrey and Pesaran [1983] and Bernanke, et. al. [1988].
10. The power of these tests, that is, the probability of correctly rejecting a false null hypothesis, has been found by Bernanke, et. al. [1988] to be quite high. Indeed, they find in their simulations that the P2-test never failed to reject a false null when the asymptotic size was 2.5 percent and when the true model had uncorrelated errors, and only failed to properly reject 2.4 percent of the time when the data were highly correlated. See their Table 4, 315.
11. Again, Barro and Rush's [1980] specification was used. That is, $\log(U/1-U)$ is the dependent variable where U is the total labor force unemployment rate, and $(G/y)_t$, the ratio of the levels of government spending and output, replaces t and G_t as explanatory variables. Eight expectational error terms and a second-order autoregressive error specification are retained.
12. The theory underlying both the Barro-Rush and the Fischer specifications implies that deviations of output or unemployment from their natural rates occur because of monetary misperceptions and, perhaps, government spending. We fit our equations using data from the 1970s, during which it is very likely that the *natural rates* of output and unemployment could not reasonably be represented by a time trend or a constant. Data we use on the natural rates of output and unemployment are the middle-expansion trend figures calculated by the Bureau of Economic Analysis (BEA) and published in the *Survey of Current Business* [Holloway 1986]. The dependent variable then becomes the difference between the logs of the actual and natural rate figures as it is these deviations that the two models considered here profess to explain. G_t and $(G/y)_t$ were retained as explanatory variables under the presumption that Barro included them to capture demand shifts rather than movements in aggregate supply.
13. We use the forecasts of Chase Econometric Associates, Inc. and its predecessors. We are grateful to Michael K. Evans, Steven J. Elgart, and Leon W. Taub for providing the forecasts used in this paper. The forecasts used were released around the third week of the second month of each quarter.

14. As the forecasts released in, say, the fourth quarter of 1979 were made available in mid-November, we define the one-quarter-ahead expected money growth rate $E_{t-4}m_t$ for $t=79:4$ as the log-difference between November's expectation of the fourth quarter's money supply and the preliminary third-quarter money supply figure, which is the most recently released measure available by the middle of November. Thus the forecasts are of a figure which will not be fully known for at least another two months, given the Federal Reserve's publication lag and the lag between the production and release of the forecasts. The aggregate being forecasted is partially known at the time of the forecast, given the weekly publication of such figures. The variable labelled $E_{t-4}m_t$ forecasts an aggregate growth rate which will be known in preliminary form five months hence, etc.
15. The forecasts used were of old M1 money supply. Thus the actual figures used were those on old M1 found in various issues of the *Federal Reserve Bulletin* and reproduced in Gordon [1987, Table A-2]. The Barro-Rush and Fischer theories do not distinguish between preliminary and revised data. We felt that the revised data best reflected the *actual* money growth which affects the economy. True *ex ante* expectations of growth, however, can only be computed based on preliminary data.
16. Mishkin [1983] and Pagan [1984] in particular provide thorough discussions of the problems associated with the two-step procedure. Whenever expectations of money growth are generated in the initial step of a two-step procedure, the validity of the proxy depends on the accuracy of the assumption of rational expectations. Our results, for both the Barro and the Fischer models, do not depend on this assumption. However, we have tested the Chase one-period-ahead money forecasts using the standard test for rationality. When the test is carried out on level forecasts, the hypothesis of rationality is not rejected at the conventional significance levels. When the test is done on the relative change forecasts, the same hypothesis is rejected at the .05 significance level but not at the .01 level.
17. Barro and Rush use quarterly growth rates while we use annual growth rates for money; thus our estimates are roughly one-fourth the size of theirs. Also, while the magnitudes of the estimates are comparable, our significance levels are reduced by our smaller sample size.
18. NA denotes that small sample significance levels are not evaluated in those cases where t -statistics obtained do not exceed their asymptotic critical values.

REFERENCES

- Barro, R. J. Unanticipated Money Growth and Unemployment in the United States. *American Economic Review*, March 1977, 101-15.
- _____. Unanticipated Money, Output, and the Price Level in the United States. *Journal of Political Economy*, August 1978, 549-80.
- Barro, R. J. and Rush, M. Unanticipated Money and Economic Activity, in *Rational Expectations and Economic Policy*, edited by S. Fischer. Chicago: University of Chicago Press, 1980, 23-48.
- Bernanke, B., Bohn, H., and Reiss, P. C. Alternative Non-Nested Specification Tests of Time-Series Investment Models. *Journal of Econometrics*, March 1988, 293-326.
- Carns, F. and Lombra, R. Rational Expectations and Short-Run Neutrality: A Reexamination of the Role of Anticipated Money Growth. *Review of Economics and Statistics*, November 1983, 639-43.
- Cornell, B. Money Supply Announcements and Interest Rates: Another View. *Journal of Business*, January 1983a, 1-24.
- _____. The Money Supply Announcements Puzzle: Review and Interpretation. *American Economic Review*, September 1983b, 644-57.
- Cox, D. R. Tests of Separate Families of Hypotheses, in *Proceedings of the Fourth Berkeley Symposium on Mathematical Statistics and Probability*. Berkeley: University of California Press, 1961, I, 105-23.

- _____. Further Results on Tests of Separate Families of Hypotheses. *Journal of the Royal Statistical Society*, 1962, Series B, 406-24.
- Davidson, R. and MacKinnon, J. G.** Several Tests for Model Specification in the Presence of Alternative Hypotheses. *Econometrica*, May 1981, 781-93.
- _____. Some Non-Nested Hypothesis Tests and the Relations Among Them. *The Review of Economic Studies*, October 1982, 51-65.
- Fischer, S.** Long-Term Contracts, Rational Expectations, and the Optimal Money Supply Rule. *Journal of Political Economy*, February 1977, 191-205.
- _____. On Activist Monetary Policy with Rational Expectations, in *Rational Expectations and Economic Policy*, edited by S. Fischer. Chicago: University of Chicago Press, 1980, 211-35.
- Gordon, R. J.** *Macroeconomics*. 4th ed. Boston: Little, Brown and Company, 1987.
- Godfrey, L. G., and Pesaran, M. H.** Tests of Non-Nested Regression Models: Small Sample Adjustments and Monte Carlo Evidence. *Journal of Econometrics*, January 1983, 133-54.
- Holloway, T. M.** The Cyclically Adjusted Federal Budget and Federal Debt: Revised and Updated Estimates. *Survey of Current Business*, March 1986, 11-17.
- Lucas, R. E., Jr.** An Equilibrium Model of the Business Cycle. *Journal of Political Economy*, December 1975, 1113-44.
- _____. Econometric Policy Evaluation: A Critique, in *The Phillips Curve and Labor Markets*, edited by K. Brunner, and A. Meltzer. Amsterdam: North-Holland, 1976, 19-46.
- McAleer, M.** Specification Tests for Separate Models: A Survey, in *Specification Analysis in the Linear Model*, edited by M. L. King, and D. E. A. Giles. London: Routledge and Kegan Paul, 1987, 146-96.
- McNees, S. K.** The Forecasting Record for the 1970s. *New England Economic Review*, September/October 1979, 33-53.
- Nichols, D. A., Small, D. H., and Webster, C. E., Jr.** Why Interest Rates Rise When an Unexpectedly Large Money Stock is Announced. *American Economic Review*, June 1983, 383-88.
- Pagan, A.** Econometric Issues in the Analysis of Regressions with Generated Regressors. *International Economic Review*, February 1984, 220-47.
- Pesaran, M. H.** On the General Problem of Model Selection. *Review of Economic Studies*, April 1974, 153-71.
- _____. A Critique of the Proposed Tests of the Natural Rate-Rational Expectations Hypothesis. *Economic Journal*, September 1982a, 529-54.
- _____. Comparison of Local Power of Alternative Tests of Non-Nested Regression Models. *Econometrica*, September 1982b, 1287-1305.
- _____. On the Policy Ineffectiveness Proposition and a Keynesian Alternative: A Rejoinder. *Economic Journal*, June 1988, 504-508.
- Phelps, E. S., and Taylor, J. B.** Stabilizing Powers of Monetary Policy under Rational Expectations. *Journal of Political Economy*, February 1977, 163-90.
- Rush, M. and Waldo, D.** On the Policy Ineffectiveness Proposition and a Keynesian Alternative. *Economic Journal*, June 1988, 498-503.
- Sargent, T. J.** Estimation of Dynamic Labor Demand Schedules Under Rational Expectations. *Journal of Political Economy*, December 1988, 1009-44.
- Taylor, J. B.** Staggered Wage Setting in a Macro Model. *American Economic Review Papers and Proceedings*, May 1979, 108-13.
- _____. Aggregate Dynamics and Staggered Contracts. *Journal of Political Economy*, February 1980, 1-23.
- Urlich, T. and Wachtel, P.** Market Response to the Weekly Money Supply Announcement in the 1970s. *Journal of Finance*, December 1981, 1063-72.