



The Effect of Monetary Changes on Interest Rates: Box-Jenkins Approach

Houston H. Stokes; Hugh Neuburger

The Review of Economics and Statistics, Vol. 61, No. 4 (Nov., 1979), 534-548.

Stable URL:

<http://links.jstor.org/sici?sici=0034-6535%28197911%2961%3A4%3C534%3ATEOMCO%3E2.0.CO%3B2-W>

The Review of Economics and Statistics is currently published by The MIT Press.

Your use of the JSTOR archive indicates your acceptance of JSTOR's Terms and Conditions of Use, available at <http://www.jstor.org/about/terms.html>. JSTOR's Terms and Conditions of Use provides, in part, that unless you have obtained prior permission, you may not download an entire issue of a journal or multiple copies of articles, and you may use content in the JSTOR archive only for your personal, non-commercial use.

Please contact the publisher regarding any further use of this work. Publisher contact information may be obtained at <http://www.jstor.org/journals/mitpress.html>.

Each copy of any part of a JSTOR transmission must contain the same copyright notice that appears on the screen or printed page of such transmission.

JSTOR is an independent not-for-profit organization dedicated to creating and preserving a digital archive of scholarly journals. For more information regarding JSTOR, please contact support@jstor.org.

THE EFFECT OF MONETARY CHANGES ON INTEREST RATES: A BOX-JENKINS APPROACH

Houston H. Stokes and Hugh Neuburger*

I. Introduction

THE work of Friedman and Schwartz (1963) has provided the impetus for much new research in the field of monetary economics. A subject upon which attention in this field has focused is monetary effects on interest rates. The difficulties that arise in identifying these effects empirically prompted us to seek to identify them in U.S. data for the period 1947 to 1978 using a variety of Box-Jenkins techniques. These techniques enable us to distinguish results showing effects that theory leads us to expect from apparently spurious results reflecting statistical problems.¹ Results obtained using several different model estimation methods have been reported

Received for publication October 6, 1976. Revision accepted for publication February 27, 1979.

* University of Illinois and Columbia University, respectively.

Computer time for this paper was provided by the Computer Center of the University of Illinois. The authors are indebted to Professors Ali Akarca, Philip Cagan, Joseph Persky, and an anonymous referee for their helpful suggestions. Any errors are the sole responsibility of the authors.

¹ In an earlier version of this paper using data on the commercial paper interest rate, M2 and real M2 for the period 1875 to 1907, we obtained similar results to those reported here. Helpful comments by a referee alerted us to potential problems of interpretation resulting from seasonality corrections of some subperiods of the money series made using a two sided filter; it is possible that these corrections altered the dynamic patterns found. Friedman and Schwartz (1963, p. 722) state that a rough seasonal correction was made in the period August 1878 to August 1881 and that no correction was made in the period June 1882 to June 1906. Private correspondence with Anna Schwartz established that data without these adjustments are not available. Further difficulties are created by various "interpolation procedures" (Friedman and Schwartz, 1970, p. 321) that were used to fill in gaps in the data. We reluctantly came to the conclusion that questions raised about the M2 series for the period 1875 to 1907 might detract from the approach we are proposing. The present paper avoids these problems because the M2 series for the period 1947II to 1978II that has been used is not seasonally adjusted. It appears that since the X-11 seasonality adjustment program incorporates a two sided filtering procedure, series seasonally adjusted with this and similar techniques should not be used to test for causality in the Granger (1969) sense. Another problem discussed in recent work of Auerbach and Rutner (1978) is the X-11 may put seasonality in the data. It seems that the degree to which the dynamic specification of macroeconomic models is an artifact of the X-11 seasonal adjustment program or other two sided adjustment techniques is a question that remains to be resolved.

including cross correlations obtained using a one filter prewhitening technique (Box and Jenkins, 1976) supplemented by a new diagnostic procedure developed in this paper, cross correlations obtained using a two filter prewhitening technique that is suitable for independence testing (Haugh, 1976; Pierce and Haugh, 1977; and Pierce, 1977) and Box-Jenkins transfer function results obtained using maximum likelihood estimation methods. These results are compared with those obtained using other methods such as cross correlations obtained without prewhitening either of the series (Friedman and Schwartz, 1976; Cagan, 1972), regressions using prewhitened data (Granger, 1973), regression coefficients obtained using raw data (Gibson, 1970b) and regression coefficients obtained from applying generalized least squares where positive and negative lags of the dependent variable are included in the regression equation (Sims, 1972). Rather than attempt to refine economic theory, we seek to show that improving statistical methods is necessary if satisfactory tests of hypotheses suggested by such theory are to be performed. For this purpose we have investigated the impact (short-run) effect of monetary aggregates on interest rates, the liquidity effect, and a more long-run effect, the income effect.²

The liquidity effect relates the rate of change in the nominal money supply to the rate of change in the interest rate or, alternatively, relates the level of the nominal money supply to the level of the interest rate.³ In analyzing the long-run effects of monetary aggregates on interest rates, it

² The effects of monetary variables, and indirectly, of price expectations on interest rates were first studied by Fisher (1896, 1907, 1930). More recent work includes that of Friedman (1961), Friedman and Schwartz (1963, 1976), Gibson and Kaufman (1968), Gibson (1970a, 1970b, 1972), Tobin (1947, 1956) and Cagan (1965, 1972).

³ Gibson (1970b) and others have noted that what is called the liquidity effect is actually one of the reduced forms of a system of simultaneous equations that assumes money is the only exogenous variable. The problem with this formulation is that the money supply may not be exogenous, since it is well known that a rise in interest rates will lead banks to operate with lower reserve requirements as they increase lending. We tested for this possibility noted by Gibson by reversing the direction of the cross correlations.

is necessary to distinguish between price or price expectations effects (Fisher, 1896, 1907, 1930) and income effects (Friedman and Schwartz, 1976; Gibson, 1970b; Cagan, 1972). The income effect arises from the stimulation given the level of income today by past increases in the money supply. The results of this stimulation are an increase in the demand for money and an increase in the interest rate in the present period.⁴ If the economy is at or near full employment when the nominal money supply is increased, prices rather than real income will rise (with a lag). The rise in prices will increase the nominal demand for money and the interest rate.⁵ Since analysis of the relationship of the nominal money stock and the interest rate would not allow us to distinguish between these two alternatives, we have also reported results of tests using real magnitudes so that this distinction can be made.

II. The Empirical Evidence

A. The Data

To insure that our results are comparable with those of Cagan and with our previous unreported results for the period 1875–1907, we have chosen M2 as our nominal monetary series, M2 deflated by the consumer price index (all items) as our real monetary series and the commercial paper interest rate as our interest series. All data have been obtained from the Monthly National Bureau of Economic Research (NBER) data bank for the period 1947II to 1978II.

B. Difficulties in Analyzing the Relationship between Two Series

One method to test whether an endogenous variable y_t is a function of present and lagged values of an endogenous variable x_t ,

$$y_t = f(x_t, x_{t-1}, \dots, x_{t-k}), \quad (1)$$

is to calculate a regression equation of the form

⁴ Cagan (1972) suggests that the lag extends six months, while Friedman (1961) and others find evidence of a longer lag. Culbertson (1960), however, is not convinced that the lag is nearly so long.

⁵ Fisher (1896, 1907, 1930) showed that the real rate of interest plus the expected change in prices is equal to the nominal rate of interest. This view suggests that as price expectations rise, the nominal rate of interest will also rise. Mundell (1963) notes, however, that the full rise in the nominal rate of interest will be damped by the fall in the real rate of interest resulting from the expected change in prices.

of (1) where y_t is regressed on present and lagged values of x_t . While Gibson and others have employed this approach to measure the relationship between interest rates and monetary aggregates, such a procedure is usually hampered by severe multicollinearity except in the unlikely case where the x variable happens to be prewhitened at the outset.⁶ Although Gibson's findings were suggestive, almost all diagnostic checking of the equations that was performed yielded very low Durbin-Watson test statistics, a sign of serial correlation of the residuals and a high degree of multicollinearity between various lags of the x variables.⁷ Because of these difficulties and serious questions concerning the interpretation of coefficient significance tests raised by Granger and Newbold (1974), the Gibson regression approach was not pursued here.

One way to avoid the difficulties of multicollinearity associated with a regression model of the Gibson type would be to attempt multiple correlation analysis (i.e., cross-correlation analysis). Although he had not obtained conclusive results when the comment was made, Friedman (1961) suggested just such an approach, noting that cross-spectral analysis offers much promise. While cross-correlation analysis does not encounter the multicollinearity problem that a regression approach does, as we show later, unless data are prewhitened cross correlation may easily yield spurious results. In the sections that follow, we discuss the problems that arise if neither of the two series is prewhitened or one of

⁶ See Gibson (1970b) and Cagan (1965) for evidence on the relationship between interest rates and monetary aggregates and Gibson (1970a) for evidence concerning interest rates and price expectations.

⁷ In the period under study the autocorrelations of nominal M2 and real M2, as measured by the modified Box-Pierce statistic (Ljung and Box, 1976) for the first 24 autocorrelations, showed very high levels of autocorrelation (7017.2 and 7850.8, respectively). Expressing these monetary aggregate terms in percentage change form caused no improvement since these statistics remained highly significant (1218.0 and 3993.1, respectively). When these four series were expressed as first differences, the modified Box-Pierce statistics for first differences of M2, real M2, percentage change in M2 and percentage change in real M2 were 3257.8, 529.35, 81.09 and 98.82, respectively. Since the modified Box-Pierce statistic is distributed as a chi-square statistic with 24 degrees of freedom, these results give some impression of the difficulty of estimating an equation of the form of (1) in the presence of such severe multicollinearity. Cagan (1972, p. 104) is well aware of this problem, noting "With so many lagged terms, the problem of collinearity among the independent variables is a matter of concern."

the series is prewhitened or both of the series are prewhitened. Our object is to show that it is possible to develop diagnostic tests for the one filter cross-correlation procedure and that the results obtained using this procedure are in many cases more meaningful than results obtained using the two filter procedure, especially if one wishes to examine the functional form of the model directly rather than to test for the independence of the series.

C. Diagnostic Checking of One Filter Cross Correlations

Box and Jenkins⁸ have studied the problems of cross correlating two data series when either or both of the series are not white noise.⁹ Their results can be summarized by four propositions, given two series x_t and y_t where both series have been filtered by the same prewhitening filter.

- If series x_t and y_t are related, and both x_t and y_t are white noise, the cross correlations of x_t on y_t and the correlations of y_t on x_t will indicate the true relationship between the series.
- If x_t and y_t are not related and if one series (say x_t) is white noise and the other series (say y_t) is not white noise, then "in this case the cross correlations have the same autocorrelation function as the process generating y_t . Thus, even though x_t and y_t are not cross correlated, the cross-correlation function can be expected to vary about zero with standard deviation $(n - k)^{1/2}$ in a systematic pattern typical of the behavior of the autocorrelation function $\rho_{yy}(j)$."¹⁰
- "If two processes are both white noise and are not cross correlated, then the covariance between the cross correlations will be zero."

⁸ See Box and Jenkins (1976), particularly p. 377.

⁹ Nelson defines white noise as "a sequence of identically and independently distributed random disturbances with mean zero and variance σ_u^2 ." See Nelson (1973, p. 31).

¹⁰ Box and Jenkins (1976, p. 377) with notational changes. If $\rho_{yy}(j)$ is the autocorrelation of series y_t , j periods apart, and $r_{xy}(k)$ is the cross correlation of series x_t and y_t , k periods apart, Box and Jenkins have proved that

$$p[r_{xy}(k), r_{xy}(k + j)] \approx p_{yy}(j)$$

where the first term refers to the autocorrelations of the cross correlations j periods apart.

- If two series are cross correlated and neither series is white noise, then it is impossible to distinguish between true cross correlations and spurious results arising from autocorrelations left in the two series.

Proposition d is especially relevant to an investigation of the problem at hand because it states that when neither series is white noise one cannot distinguish between spurious and non-spurious cross correlations.¹¹ Hence cross correlations between lagged monetary variables and interest rates reported by Cagan (1972), Friedman and Schwartz (1976) and others raise serious problems of interpretation. As we will show later empirically, if there is a relationship between two series lagged k periods and if the series that are cross correlated are themselves autocorrelated, there will be relatively large cross correlations plus or minus a number of periods about the k period lag and there may be hard-to-identify spurious cross-correlation spikes at other lags as well.

Proposition b, upon which we base our diagnostic checking procedure, states that assuming x_t is a white noise series (i.e., prewhitened) and y_t is not a white noise series, if there is no relationship between x_t and y_t , then the autocorrelations of the cross correlations will be similar to the autocorrelations of the not completely prewhitened series y_t . Leaving to be explained in the next section why the cross correlations obtained using the one filter procedure will approximate the impulse response function between the two series (Box and Jenkins, 1976), we outline our proposed diagnostic procedure based on the above four propositions.

- First, an autoregressive integrated moving average (ARIMA) model is identified and estimated to reduce the original x series to white noise (as measured by the Ljung-Box (1976) modified Q -statistic).¹²

¹¹ Proposition d indicates the difficulty of interpreting cross correlations of unfiltered data. Even if the autocorrelations of the two series were known, it would not be much help because the cross correlations that these autocorrelations generate would depend upon complicated interaction effects. See Box and Jenkins (1976).

¹² To test whether a series is prewhitened, we have used the modified Q -statistic, which is distributed as a chi-square statistic with $K - R$ degrees of freedom where K is the number of autocorrelations in the statistic and R is the num-

2. Next, the x series filter is used to "filter" the y series and the autocorrelations of this new y series are calculated (the x series filter, however, will usually not prewhiten the y series entirely, in which case the diagnostic checks described in step 4 below are needed).
3. Two sets of cross correlations are calculated, one relating filtered y_t to present and past values of prewhitened x_t and one relating prewhitened x_t to present and past values of filtered y_t .
4. The two vectors of cross correlations are then individually autocorrelated and these two sets of autocorrelations are compared separately to the autocorrelations of the not completely prewhitened y_t series. A diagnostic test statistic can be calculated to test whether the two vectors of autocorrelations are the same.¹³

ber of parameters estimated in the filter (if the series is filtered). The modified Q -statistic, which is defined as

$$MQ(k) = \sum_{i=1}^k (r_i^2 / (N - i)) (N(N + 2))$$

where N is the number of observations in the series being autocorrelated and r_i is the i^{th} autocorrelation, is discussed in Ljung-Box (1976, p. 26).

¹³ The autocorrelations can be compared via inspection of the position and significance of the spikes or by the use of a test statistic. If r_i and a_i are corresponding elements of two autocorrelation functions, the statistic

$$h(i) = (a_i - r_i)^2 / (v_a + v_r)$$

is distributed as a chi-square statistic of degrees of freedom 1 where v_a and v_r are the variance of a_i and r_i . Since the sum of two chi-square statistics with degrees of freedom f_1 and f_2 has a chi-square distribution with degrees of freedom $f_1 + f_2$ (Brownlee, 1965, p. 84), it is possible to test k pairs of autocorrelations at one time with the chi-square statistic $H(k)$ where

$$H(k) = \sum_{i=1}^k h(i).$$

The variance of each element of the autocorrelation function has been calculated using Bartlett's formula (Box and Jenkins, 1976, p. 177)

$$v_a = (1/N_a) \left(1.0 + 2 \left(\sum_{k=1}^i a_k^2 \right) \right)$$

where N_a = number of data points in series a . Work on refinements of the H -statistic is currently under way. An unresolved question is whether one should test all autocorrelation coefficients, only those autocorrelation coefficients that are significant to some specified level or some other subset such as individual elements. Assuming an appropriately constructed H -statistic, a significant H -statistic indicates that the two vectors of autocorrelations are not similar.

If the autocorrelations of the not completely prewhitened y series are similar to the autocorrelations of either of the vectors of cross correlations, then proposition b requires that we refrain from giving cross correlations in that vector an economic interpretation. In summary, in order for us to argue that x_{t-k} is significantly correlated with y_t , two conditions must be met. First, there must be a spike in the cross correlation vector at lag k at least equal to two standard deviations and second, the autocorrelations of the not completely prewhitened series (y_t) must *not* be similar to the autocorrelations of the cross correlations of filtered y_t on prewhitened x_t . In our view as long as proper diagnostic tests are used (and these tests can also be used in transfer function estimation),¹⁴ the one filter cross-correlation procedure is a suitable way to search for the dynamics of a relationship (or the lack of a relationship) between any two series. To support this view we present in a later section results obtained using this technique and results obtained using a transfer function model between x_t and y_t so that a comparison between them can be made.

D. Problems of Independence Testing

The seminal article by Granger (1969) concerning the necessary and sufficient conditions for the various causal relationships that may exist between x and y led to the development of a number of statistical procedures intended to test Granger causality formally. In this section we relate these

¹⁴ The diagnostic testing procedure described above has applications both to the interpretation of one filter cross correlations and to transfer functions. Consider the following problem in transfer function estimation. The experimenter has made a preliminary determination of the form of the transfer function using either the one or two filter method (Box and Jenkins, 1976, ch. 11 and Granger and Newbold, 1977, ch. 7) and notes after estimation that there are spikes in the autocorrelation of the residual, suggesting that the noise model may have to be changed, and that there are also spikes in the cross correlations of the residual and the prewhitened input, suggesting that the input model formulation may have to be changed. What makes such a situation confusing is that since the residual is not clean (prewhitened) it is possible that spikes seen in the cross correlations are spurious, having occurred via the mechanism described in proposition b. Unless the possibility that these cross correlation spikes are spurious is tested using the diagnostic procedure described above, the experimenter may try to remove them by changing the transfer function input specification when they are actually spurious and would disappear if the noise model were correct. Testing with the suggested diagnostic procedure would streamline the transfer function estimation process and reduce the number of false tries to remove "spikes."

procedures to the one filter technique and discuss some difficulties that have arisen concerning them.

In a very provocative article, Sims (1972) suggested that a test for Granger causality could be performed if one calculated an OLS regression on data filtered with the *same filter* where future, present and past values of x are regressed on y and the filter has been chosen to prewhiten the error term. According to Sims, significance on the future value of x would suggest "feedback" of y on x and would imply that x is not exogenous, while significance on lagged values of x would suggest that x causes y . Recently, however, Pierce and Haugh (1977) showed there is a substantial possibility that this procedure will indicate causality where none exists because Sims corrected only for serial correlation of the residuals and the right hand variables in his regression would most likely not be prewhitened by the same filter that was selected to prewhiten the residual.¹⁵ Another weakness of the Sims procedure for our purposes is that since it is a test for independence, the presence of future terms on the right hand side of the equation makes it hard to interpret Sims' results in an economic framework, even if we ignore the estimation problems alluded to above. In view of these difficulties and the susceptibility of the Sims procedure to many of the problems of the usual regression approach, we have not pursued it further. The test for independence that we do report and which we relate to the one filter model identification procedure and the transfer function approach is the Haugh (1976), Pierce and Haugh (1977), Pierce (1977) two filter test for independence.

Given that x and y are two economic time series, a Box-Jenkins univariate¹⁶ prewhitening filter can be written as

¹⁵ Pierce and Haugh (1977, p. 279) note that Sims (1972) selected his filter so that the residuals of his equation were uncorrelated. They note, "One suspects that in the application of this procedure, in those cases where the prechosen filter leaves substantial serial correlation in the filtered series, it is possible that causality may be believed to have been found where it does not exist. That this may have occurred in the case of money and income, the principal focus of this study by Sims (1972), is suggested by the findings of Feige and Pearce (1976)." Additional discussion of this point is given in Feige and Pearce (1979).

¹⁶ For a discussion of Box-Jenkins prewhitening filters, see Box and Jenkins (1976) and Nelson (1973).

$$U_{xt} = \frac{\lambda_x(B)}{\theta_x(B)} x_t \quad (2)$$

and

$$U_{yt} = \frac{\lambda_y(B)}{\theta_y(B)} y_t \quad (3)$$

where U_{xt} and U_{yt} are the resulting prewhitened series and $\lambda_x(B)/\theta_x(B)$ and $\lambda_y(B)/\theta_y(B)$ are the ratios of autoregressive (AR) and moving average (MA) polynomials in B of order p and q .¹⁷ If the two prewhitened series, U_{xt} and U_{yt} , are cross correlated it is possible to identify the relationship

$$U_{yt} = \frac{\omega'(B)}{\delta'(B)} U_{xt} + e'_t \quad (4)$$

from which the basic relationship between the two raw series can be derived by substituting in for the two filtered series¹⁸

$$y_t = \frac{\theta_y(B)}{\lambda_y(B)} \frac{\lambda_x(B)}{\theta_x(B)} \frac{\omega'(B)}{\delta'(B)} x_t + e_t. \quad (5)$$

While one advantage of cross correlating two prewhitened series to identify $\omega'(B)/\delta'(B)$ is that the complete prewhitening of both series leaves no possibility of spurious cross correlations, there are two serious problems with this technique. The first, discussed by Sims (1977), is that cross correlations obtained by the two filter procedure are biased in favor of *not* finding a relationship between x and y , while the second, and related problem is that the Haugh (1976) S -statistic (or modified S -statistic) does not always have an unambiguous interpretation when used in causality tests. We will discuss the nature of these problems in detail below.

Sims' argument is that since $\omega'(B)/\delta'(B)$ is estimated conditional on the estimation of the prewhitening filters $\theta_y(B)/\lambda_y(B)$ and $\theta_x(B)/\lambda_x(B)$ only in the case where all other variables affecting Y are not correlated with X will $\omega'(B)/\delta'(B)$ be an unbiased estimate of the true relationship.

If, instead of using two filters, the same filter (i.e., the one filter technique) were used for both series, then $\theta_y(B) = \theta_x(B)$ and $\lambda_y(B) = \lambda_x(B)$. The resulting relationship found between the two

¹⁷ B is the lag operator defined such that $BY_t = Y_{t-1}$. In general $p_y \neq p_x$ and $q_y \neq q_x$.

¹⁸ We have ignored the specification of the noise model in this formulation. In terms of our notation $e_t = \theta_y(B)/\lambda_y(B) e'_t$.

series $\omega'(B)/\delta'(B)$ in this case will *not* be biased as it might have been had the two filter procedure been used, since the prewhitening filters cancel out and estimation is no longer conditional as with the two filter procedure.¹⁹ An additional advantage is that this relationship can be interpreted directly.

The bias of the two filter procedure is downward. Therefore if lack of independence is found it is possible to assert at least for the given confidence level that the series are not independent because our test is biased against such a finding.²⁰ The downward bias of the test makes it impossible to conclude that the two series are independent because no matter how small the cross correlations are, it is not possible to accept the null hypothesis of independence.

The second and related problem with the two filter procedure concerns the interpretation of the suggested test statistic to be used for the test for independence. Haugh (1976) proposed the use of the modified *S*-statistic (*MS*), distributed as a chi-square with $2M+1$ degrees of freedom:

$$MS = N^2 \sum_{k=-M}^M \frac{r_{12}(k)^2}{(N - |k|)} \quad (6)$$

where N is the number of observations used to form the cross correlations, M is the number of lags in each cross-correlation vector and $r_{12}(k)$ is the cross correlation between series 1 and series 2 lagged k periods. Instead of using the Haugh statistic on both vectors of the cross correlations at once to test for independence as suggested by Haugh, Pierce (1977) calculated his statistic on one side only (using only one cross-correlation vector) to test for the various causal relationships defined by Granger (1969).

Such a procedure has been criticized on the ground that the distribution of $r_{12}(k)$ depends on non-zero population cross correlations even when such non-zero population cross correlations occur only at lags different from the sample cross correlations used in the calculation of the

modified *S*-statistic. If this problem is present, the standard errors of the sample cross correlations may be inflated (Pierce and Haugh, 1977). The implication of this second problem is that applying the modified *S*-statistic in one sided tests for causality using the two filter method may yield results whose interpretation is not clear. In our results for the two filter method, listed in table 4, we have reported tests for independence which must be considered bearing the above limitation in mind, and we have reported tests for one way causality and feedback only for illustrative purposes and for comparison with results obtained using transfer function methods.

Employing a similar technique, Granger (1973) attempted tests for causality using regression models that contained data prewhitened with ARIMA models such as (2) and (3). Because such a procedure suffers from exactly the same features and limitations as the Haugh two filter cross correlation procedure, we have not explored it here.²¹

Pierce and Haugh (1977) propose as a solution the estimation, using maximum likelihood techniques (Box-Jenkins, 1976), of a transfer function model where "there do exist distributional results for auto and cross correlations of the residuals of fitted (one-sided) dynamic regression models of the form:

$$Y_t = \tau(B)x_t + \xi(B)e_t \quad (7)$$

where e_t is white noise independent of x_t ." If this procedure is used, the appropriate test of the correct dynamic functional form of the model $\tau(B)$ can be performed (if the noise model is correct) by cross correlating the prewhitened residual with the prewhitened input U_{xt} (obtained from equation (2)).²²

²¹ For a detailed discussion of these topics as well as causality testing in general, see Pierce and Haugh (1977, pp. 283–287).

²² Pierce and Haugh (1977, p. 284) note, "The asymptotic distribution of these residual cross correlations is known (Haugh (1972), Pierce (1972)), and is such that only a degree-of-freedom correction need be made in the χ^2 statistic." The degrees of freedom for the modified *S*-statistic calculated for M terms relating the residual to lags of the prewhitened input is $M - q$ where q is the number of parameters estimated in the transfer function input model. The degrees of freedom for the same statistic calculated for M terms relating the prewhitened input to lags of the residual (to test for feedback) is M .

¹⁹ By using this procedure, we have not "added" anything to the relationship between the two series and have taken out only the trends and seasonality common to both.

²⁰ Defining omitted variables as having the form $c(B)x_t$, Sims (1977) notes, "Anyone versed in the theory of least-squares regression will recognize this as involving a bias in favor of the null hypothesis, except in the special case when the omitted variables are uncorrelated with the included variables."

E. The Transfer Function Approach

Equation (7) is equivalent to a model of the form of (1) with the addition of a noise model $\xi(B)$. Equation (7) can be estimated²³ using a more parsimonious parameterization if we express $\tau(B)$ and $\xi(B)$ as ratios of two polynomials in B as shown in (8):

$$y_t = \frac{\omega(B)}{\delta(B)} x_t + \frac{\theta(B)}{\phi(B)} e_t. \quad (8)$$

Assuming y_t is the commercial paper interest rate and x_t is the nominal money supply (M2), it is possible to reparameterize equation (8) as

$$y_t = \frac{\omega^*(B)}{\delta^*(B)} x_t + \frac{\omega^{**}(B)}{\delta^{**}(B)} (1 - B)x_t + \frac{\theta(B)}{\phi(B)} e_t. \quad (9)$$

Note that equation (9) also requires the assumption that changes in the money supply $((1 - B)x_t)$ are sufficiently unrelated to the level of the money supply (x_t) to identify each distinct term.²⁴ In this case $\omega^*(B)/\delta^*(B)$ may be said to represent the liquidity effect and $\omega^{**}(B)/\delta^{**}(B)$ the income effect.²⁵ Because it is difficult to estimate the functional form of equation (9), we may wish to use the functional form of (8) so that the estimated ratio of polynomials $\omega(B)/\delta(B)$ is

$$\frac{\omega(B)}{\delta(B)} = \frac{\omega^*(B) \delta^{**}(B) + \delta^*(B) \omega^{**}(B)(1 - B)}{\delta^*(B) \delta^{**}(B)} \quad (10)$$

or is a combination of both the income and the liquidity effects.

Because a transfer function model of the form of (8) captures these two effects *only* in combination as shown above, an important difference between this method and the one filter technique should be clear. When the one filter technique is

used to estimate cross correlations between y_t and x_t and between y_t and $(1 - B)x_t$, it is possible to obtain *separate* measures of each of the two effects (depending on the degree of collinearity between the level of money and the first difference of money). In summary, we have shown how the one filter cross-correlation technique may be used to estimate the discrete effects that are captured together in a transfer function of the form of equation (8). As noted earlier, however, the transfer function is the most appropriate technique to test for independence and/or causality. For this reason a comparison of our transfer function results with those obtained using the two filter Haugh-Pierce independence testing method and the one filter cross-correlation procedure (with diagnostic testing) is advisable.

Our final step is to show how findings from the one filter cross-correlation procedure, transfer function procedure and regression approach are related. Although there are statistical advantages to estimating a model like (1) as a ratio of polynomials (as in (8)) to make such results directly comparable to the lag structure obtained by the regression approach of Gibson, Cagan and others, we must express $\omega(B)/\delta(B)$ as follows:²⁶

$$\omega(B)/\delta(B) = V(B) = v_0 + v_1B + v_2B^2 + v_3B^3 + \dots + v_kB^k. \quad (11)$$

If, for example, x_t influences y_t with a lag of 16 months and x_t is autocorrelated, results obtained with the one filter cross correlation procedure will show a spike only at the 16th period while, because of the autocorrelation of the x_t series, the impulse response weights ($V(B)$) will show a distribution of higher values around the 16th lag that will be similar to the regression coefficients obtained by Gibson.²⁷

F. The Results

Prewhitening ARIMA models (reported in table 1) have been estimated using real M2, nominal M2, the commercial paper interest rate and percentage changes of these variables. Inspection of the individual autocorrelations of the re-

²³ Methods of transfer function model identification are not discussed in this paper. One filter procedures are discussed in Box and Jenkins (1976, chs. 10 and 11); two filter procedures are discussed in Box and Haugh (1977) and Granger and Newbold (1977).

²⁴ In the period investigated here the correlation between M2 and $(1 - B)M2$ is 0.79979 and the correlation between $M2/P$ and $(M2/P)(1 - B)$ is 0.21819. Unless the two inputs in equation (9) are themselves unrelated, the order in which the two input models are identified will influence their functional form but not the predictive power of the equation.

²⁵ All other influences on the interest rate are captured in the noise model.

²⁶ $V(B)$ is equivalent to $\tau(B)$ in equation (7) and is the input response weights vector. See Box and Jenkins (1976, chs. 10 and 11).

²⁷ We appreciate the advice of Professor Philip Cagan concerning the need for a discussion of the different approaches to the estimation problem.

TABLE 1.—ESTIMATED MONTHLY PREWHITENING MODELS, 1947II–1978II

<u>Real M2 Not Seasonally Adjusted</u>						
$(1 - B)^2(M2/P)_t = (1 - .52399B - .11366B^2 - .08189B^7 - .11434B^{25})(1 - .14935B^{12}) u_t$						
(10.08)	(2.19)	(2.00)	(2.77)	(2.76)		
RSS = .055586	RSE = .012341	MQ(24) = 18.974				
<u>Nominal M2 Not Seasonally Adjusted</u>						
$(1 + .46005B + .45503B^2)(1 - B)^2M2_t = (1 - .16193B^3 - .13687B^4 - .29612B^7) u_t$						
(-8.77)	(-8.29)	(2.74)	(2.61)	(5.66)		
RSS = 429.23	RSE = 1.0874	MQ(24) = 23.930				
<u>Commercial Paper Interest Rate 4-6 Months Not Seasonally Adjusted</u>						
$(1 - .54743B)(1 - .27904B^{12})(1 - B)CPI_t = (1 - .14648B^2 - .23717B^6 - .15967B^7 - .25687B^{14}) u_t$						
(11.08)	(5.22)	(2.81)	(4.78)	(3.28)	(5.06)	
RSS = 20.469	RSE = .24115	MQ(24) = 26.824				
<u>Percentage Change in Real M2 Not Seasonally Adjusted</u>						
$(1 - B)PCRM2_t = (1 - .58947B - .10965B^2 - .14717B^5) u_t$						
(11.51)	(2.03)	(3.51)				
RSS = .0065424	RSE = .0042164	MQ(24) = 28.093				
<u>Percentage Change in Nominal M2 Not Seasonally Adjusted</u>						
$(1 - B)PCM2_t = (1 - .52569B - .15714B^2 - .21841B^7) u_t$						
(10.53)	(3.14)	(5.80)				
RSS = .0024893	RSE = .0026008	MQ(24) = 17.234				
<u>Percentage Change Commercial Paper Interest Rate 4-6 Months Not Seasonally Adjusted</u>						
$PCCPI = (1 + .51849B - .13889B^6 - .15971B^7 - .13576B^{14} - .11325B^{18} - .20302B^{19} - .14113B^{20} - .13730B^{22}) u_t + .0069969$						
(-12.01)	(2.77)	(3.18)	(3.11)	(2.22)	(3.53)	(2.59) (2.92) (5.00)
RSS = .93129	RSE = .050651	MO(24) = 22.693				

Note: For data sources see text. *t*-statistics listed under coefficients. MQ(24) is the modified *Q*-statistic (Ljung and Box, 1976) for the first 24 autocorrelations of the residual. The appropriate degrees of freedom for MQ(24) is 24 - *k* where *k* is the number of estimated coefficients in the model. RSS is the residual sum of squares and RSE is the residual standard error. All models have been estimated using techniques suggested by Box and Jenkins (1976).

siduals as well as the modified Box-Pierce statistics indicates that these filters reduce their respective series to white noise.

Evidence for the liquidity effect obtained by applying the one filter cross-correlation technique to both the real and the nominal money stock is reported in table 2. As theory predicts, this evidence indicates a significant negative correlation in the zero period (of -0.125 and -0.130, respectively). Inspection of the diagnostic tests suggests that the only other significant spike at lag 7 (between the percentage change in the commercial paper interest rate and lags of the percentage change in real balances) is spurious.²⁸

Evidence for the income effect is given in table 3. In the case of the nominal money supply, we find a spike at lag 16 (also found later in the

transfer function results) which is consistent with the suggestion of Friedman (1961) that the effect of a change in the nominal money supply is to increase interest rates 16 months later. This finding is confirmed by diagnostic checking. The feedback of interest rates on the nominal money stock, which indicates that a rise in the interest rate will slow the rate of change of the nominal money stock several periods later, cannot be rejected by diagnostic tests although it is not of the expected sign. Perhaps it reflects a switch out of M2 and into other monetary assets.

Evidence for the income effect using real balances indicates a slightly longer lag structure than that found using nominal balances.²⁹ The

²⁸ The statistic *h*(7), which tests whether there is a significant difference between the autocorrelations -0.15 and -0.18, was found to be 0.02769, which is not significant. The value of this statistic suggests that the spike at lag 7 (0.115) is most likely an echo of other information in the vector (possibly the significantly negative coefficient -0.125 at lag 0).

²⁹ One difference observed is that for real balances there appear to be significant correlations between the level of interest rates and lags in the change in real balances at 0 to 3, while in the case of nominal balances there is a significant correlation only at lag 0. The very low values of the statistics *h*(1), *h*(2) and *h*(3) (0.08247, 0.002155, and 0.02623, respectively) suggest that the spikes at lags 1 to 3 are spurious. Visual inspection of the two vectors confirms this view.

TABLE 2.—TEST OF LIQUIDITY EFFECT, 1947II–1978II

Series 1—Percentage Change in Real Balances												
Series 2—Percentage Change in Commercial Paper Interest Rate 4–6 Months												
Cross Correlations of Data Filtered with Series 1's Filter												
Series 2 on Lags of Series 1						Series 1 on Lags of Series 2						
Lag	0	-.125				Lag	0	-.125				
	1	-.003					1	-.046				
	2	.029					2	-.047				
	3	.023					3	-.073				
	4	.074					4	-.040				
	5	-.001					5	-.060				
	6	.050					6	-.057				
	7	.115					7	-.021				
	8	.031					8	-.016				
	9	-.034					9	-.044				
	10	.029					10	.034				
	11	.001					11	.012				
	12	.036					12	-.068				
	13	.055					13	-.072				
	14	.074					14	.042				
	15	.006					15	-.046				
	16	.069					16	.068				
	17	-.007					17	.046				
	18	-.065					18	-.074				
	19	.103					19	.022				
	20	.025					20	-.042				
	21	-.017					21	-.036				
	22	-.011					22	.058				
	23	.072					23	.008				
	24	.060					24	.040				
						SE = .051917						
Diagnostic Checking												
Autocorrelations of Series 2 Which Has Been Filtered by the Series 1 Filter Lag												
Lag												
1–12	.19	-.14	-.01	-.05	-.05	-.09	-.15	-.06	.05	.00	.04	.11
13–24	-.02	-.07	.05	-.03	-.06	-.08	-.14	-.04	.01	-.11	.06	.11 $MQ(24) = 67.39$
Autocorrelations of Cross Correlations of Series 2 on Lags of Series 1												
Lag												
1–12	.17	-.28	-.17	.01	.24	.32	-.18	-.19	.26	.21	.07	-.02
13–24	-.19	-.23	.30	.11	.02	-.07	-.20	-.06	.18	-.05	-.16	-.06 $MQ(24) = 59.15$
Autocorrelations of Cross Correlations of Series 1 on Lags of Series 2												
Lag												
1–12	.20	.07	.17	-.18	.15	.22	.03	.12	.16	.10	-.05	.08
13–24	-.01	-.21	.18	-.03	-.10	.19	-.18	-.01	.09	-.16	-.10	-.11 $MQ(24) = 33.24$

Series 1—Percentage Change in Nominal Balances

Series 2—Percentage Change in Commercial Paper Interest Rate 4–6 Months

Cross Correlations of Data Filtered with Series 1's Filter

Series 2 on Lags of Series 1						Series 1 on Lags of Series 2					
Lag	0	-.130				Lag	0	-.130			
	1	.047					1	-.062			
	2	.060					2	-.093			
	3	.098					3	-.093			
	4	.097					4	-.043			
	5	.018					5	.017			
	6	.093					6	-.046			
	7	.033					7	.078			
	8	-.013					8	-.023			
	9	-.082					9	-.060			
	10	.014					10	.092			
	11	.039					11	-.055			
	12	-.009					12	-.076			
	13	.059					13	-.044			
	14	.055					14	.085			
	15	-.003					15	-.028			

TABLE 2.—(Continued)

16	.104	16	.047
17	-.061	17	.007
18	-.029	18	-.071
19	.049	19	.037
20	-.006	20	-.075
21	-.045	21	-.074
22	-.040	22	.134
23	.106	23	.081
24	-.097	24	.031

SE = .051917

Diagnostic Checking

Autocorrelations of Series 2 Which Has Been Filtered by the Series 1 Filter

Lag

1-12	.15	-.14	-.02	-.08	-.22	-.20	-.01	.05	.09	.02	.04	.09
13-24	-.04	-.08	.04	.01	-.03	-.07	-.11	-.03	-.02	-.12	.08	.12 $MQ(24) = 86.37$

Autocorrelations of Cross Correlations of Series 2 on Lags of Series 1

Lag

1-12	.04	-.03	-.04	-.06	-.12	-.04	.10	-.04	.13	.11	.02	.08
13-24	.05	-.12	.12	-.09	.11	-.17	-.08	-.06	.03	.11	.00	.04 $MQ(24) = 14.18$

Autocorrelations of Cross Correlations of Series 1 on Lags of Series 1

Lag

1-12	.12	-.03	.08	-.26	-.17	.04	-.07	.10	.42	-.02	-.10	.14
13-24	-.18	-.20	.12	-.10	.05	.25	-.04	-.03	.04	-.26	-.11	-.04 $MQ(24) = 49.69$

Note: For a discussion of data sources and the methods of analysis, see text. Only 25 of 55 cross correlations have been reported. In the calculation of the autocorrelations of the cross correlations the complete set of 55 has been used. For a discussion of the filters used, see table 1 and text.

TABLE 3.—TEST OF INCOME EFFECT 1947II-1978II

Series 1—Change of Real Balances Series 2—Level of Commercial Paper Interest Rate 4-6 Months Cross Correlations of Data Filtered with Series 1's Filter												
Series 2 on Lags of Series 1						Series 1 on Lags of Series 2						
Lag	0	-.232				Lag	0	-.232				
	1	-.213					1	-.153				
	2	-.136					2	-.136				
	3	-.113					3	-.078				
	4	-.054					4	.012				
	5	-.037					5	.041				
	6	-.035					6	.050				
	7	.015					7	.035				
	8	.030					8	.013				
	9	-.052					9	.011				
	10	-.052					10	-.012				
	11	-.051					11	-.070				
	12	-.028					12	-.060				
	13	-.029					13	.006				
	14	.023					14	.036				
	15	.049					15	.024				
	16	.073					16	.065				
	17	.107					17	.035				
	18	.061					18	.023				
	19	.121					19	.043				
	20	.129					20	-.019				
	21	.052					21	.007				
	22	.045					22	.006				
	23	.054					23	-.045				
	24	.027					24	-.025				

SE = .051988

Diagnostic Checking

Autocorrelations of Series 2 Which Has Been Filtered by the Series 1 Filter

Lag												
1-12	.81	.55	.36	.15	-.06	-.20	-.18	-.07	.03	.13	.25	.32
13-24	.22	.08	-.02	-.17	-.30	-.40	-.42	-.34	-.23	-.12	.02	.11 $MQ(24) = 794.99$

TABLE 3.—(Continued)

Autocorrelations of Cross Correlations of Series 2 on Lags of Series 1													
Lag													
1-12	.77	.54	.40	.24	.14	.07	-.02	.00	.09	.11	.12	.08	
13-24	-.01	-.08	-.09	-.17	-.22	-.25	-.32	-.25	-.12	-.08	-.02	.05	$MQ(24) = 100.11$
Autocorrelations of Cross Correlations of Series 1 on Lags of Series 2													
Lag													
1-12	.65	.33	.05	-.22	-.32	-.33	-.22	-.02	.15	.25	.31	.25	
13-24	.06	-.09	-.15	-.26	-.23	-.19	-.18	.01	.12	.19	.23	.15	$MQ(24) = 103.09$
Series 1—Change in Nominal Balances Series 2—Level of Commercial Paper Interest Rate 4-6 Months													
Cross Correlations of Data Filtered with Series 1's Filter													
Series 2 on Lags of Series 1							Series 1 on Lags of Series 2						
Lag	0	-.169					Lag	0	-.169				
	1	-.094						1	-.127				
	2	-.073						2	-.126				
	3	-.014						3	-.088				
	4	.022						4	-.033				
	5	-.010						5	.034				
	6	-.013						6	.031				
	7	-.021						7	.028				
	8	-.005						8	-.044				
	9	-.063						9	.019				
	10	-.067						10	.012				
	11	-.018						11	-.051				
	12	.005						12	-.052				
	13	.045						13	-.043				
	14	.048						14	.025				
	15	.071						15	-.054				
	16	.137						16	-.021				
	17	.071						17	-.048				
	18	.080						18	-.074				
	19	.086						19	-.063				
	20	.084						20	-.118				
	21	.036						21	-.035				
	22	.003						22	.015				
	23	.077						23	-.075				
	24	.021						24	-.083				
							SE = .052129						
Diagnostic Checking													
Autocorrelations of Series 2 Which Has Been Filtered by the Series 1 Filter													
Lag													
1-12	.84	.67	.54	.42	.28	.17	.18	.17	.16	.16	.18	.16	
13-24	.06	-.02	-.08	-.17	-.25	-.32	-.36	-.36	-.34	-.30	-.26	-.24	$MQ(24) = 1018.2$
Autocorrelations of Cross Correlations of Series 2 on Lags of Series 1													
Lag													
1-12	.72	.52	.42	.35	.27	.18	.18	.10	.07	.11	.08	-.03	
13-24	-.14	-.13	-.09	-.18	-.16	-.20	-.23	-.24	-.15	-.08	-.15	-.13	$MQ(24) = 110.55$
Autocorrelations of Cross Correlations of Series 1 on Lags of Series 2													
Lag													
1-12	.72	.55	.45	.41	.35	.24	.28	.33	.27	.24	.22	.23	
13-24	.13	.01	.09	.07	.09	.00	-.07	-.02	-.03	-.04	-.04	-.08	$MQ(24) = 129.71$

Note: For a discussion of data sources and the methods of analysis, see text. Only 25 of 55 cross correlations have been reported. In the calculation of the autocorrelations of the cross correlations the complete set of 55 has been used. For a discussion of the filters used, see table 1 and text.

same negative feedback found with nominal balances is found with real balances and may reflect the way in which a rise in expected prices causes (i.e., Fisher effect) increases in the interest rate with subsequent increases in prices lowering real balances.

The results of the Haugh-Pierce two filter test for independence for various specifications of the monetary aggregates and interest rates are given in table 4. Except for the modified *S*-statistic for the percentage change in real M2 and the percentage change in the commercial paper interest

TABLE 4.—HAUGH-PIERCE χ^2 TEST FOR INDEPENDENCE
(both series filtered by their respective filters from table 1)

	<i>CPI - M2/P</i>	<i>CPI - M2</i>	<i>CPI - PCRM2</i>	<i>CPI - PCM2</i>	<i>PCCPI - PCRM2</i>	<i>PCCPI - PCM2</i>
Causality plus						
Instantaneous						
Causality						
DF 25	43.12 ^a	46.12 ^a	37.00 ^a	41.63 ^a	32.72	24.01
DF 37	58.62 ^a	74.92 ^a	53.79 ^a	69.36 ^a	46.18	51.79
DF 55	96.88 ^a	110.47 ^a	86.42 ^a	109.44 ^a	67.99	73.81 ^a
Feedback						
DF 24	42.53 ^a	66.07 ^a	33.05	42.50 ^a	29.30	43.50 ^a
DF 36	56.36 ^a	89.84 ^a	42.98	53.73 ^a	35.45	52.77 ^a
DF 54	63.31	114.90 ^a	46.33	62.47	41.62	60.96
Independence						
DF 49	85.66 ^a	112.19 ^a	70.05 ^a	84.12 ^a	62.01	67.50 ^a
DF 73	114.98 ^a	164.77 ^a	96.77 ^a	123.09 ^a	81.63	104.56 ^a
DF 109	160.18 ^a	225.37 ^a	132.75 ^a	171.92 ^a	109.61	134.77 ^a

Note: *CPI* = commercial paper interest rate.*PCCPI* = percent change commercial paper interest rate.*M2/P* = Real M2.*M2* = Nominal M2.*PCRM2* = percentage change real M2.*PCM2* = percentage change M2.

All statistics reported are the Haugh modified *S*-statistic (Haugh, 1976, p. 383). The exact distribution of the *S*-statistic is known only for tests for independence (Pierce and Haugh, 1977, p. 284). Pierce (1977, p. 15) indicated that causality tests and feedback tests using the two filter method and the *S*-statistic underestimate significance if some significant cross correlations are not in the test statistic. The reported results must be interpreted with this limitation in mind. The independence tests can be interpreted directly, although the conditional estimation bias discussed in the text still applies. For further discussion of the results, see the text.

^a Significant at the 95% level or better.TABLE 5.—TRANSFER FUNCTIONS RELATING THE COMMERCIAL PAPER INTEREST RATE TO REAL M2 AND
NOMINAL M2 IN THE PERIOD 1947II–1978II
Real M2 Equation

$$(1 - B)CPI_t = (-2.6673 + 2.6595B^7 - 1.3607B^9 + 2.8204B^{14} + 2.2869B^{19})(1 - B)^2(1 - .95536B)^{-1}M2/P_t + (1 - .13791B^2 - .32174B^6 - .25356B^7)(1 + .28915B^{12})(1 - .49543B)^{-1}e_t$$

(2.80) (2.62) (1.43) (-3.00) (2.49) (46.02) (2.57)

(6.08) (5.14) (-5.03) (9.53)

RSS = 19.475 RSE = .23968

Residual Diagnostic Check $MQ(24) = 26.19$ (no significant spikes remaining) $MSL(37) = 39.36$ $MSR(36) = 52.12$ $MSB(73) = 91.483$

No significant spikes were observed between cross correlations of the residual and lags of prewhitened (*M2/P*). Three significant spikes were observed for cross correlations between prewhitened *M2* and lags of the residual at 2, 3 and 11 of -0.149, -0.124 and -0.111, respectively.

Lag	Model Implied Impulse Response Weights	Cumulative Effect
0	-2.667	-2.667
1	-2.548	-5.215
2	-2.435	-7.650
3	-2.326	-9.976
4	-2.222	-12.198
5	-2.123	-14.321
6	-2.028	-16.349
7	0.722	-15.627
8	0.690	-14.937
9	-0.702	-15.639
10	-0.670	-16.309
11	-0.640	-16.949
12	-0.612	-17.561
13	-0.585	-18.146
14	2.262	-15.884
15	2.161	-13.723
16	2.064	-11.659
17	1.972	-9.687
18	1.884	-7.803
19	4.087	-3.716

TABLE 5.—(Continued)

20	3.905	0.189
21	3.730	3.919
22	3.564	7.483
23	3.405	10.888
24	3.253	14.141

Nominal M2 Equation

$$(1 - B)CPI_t = (-.029055 + .040543B^3 + .04010B^4 - .030707B^6 + .031577B^{16} + .010846B^{23}) * (1 - .62873B)^{-1}$$

(−4.38) (−4.14) (−3.89) (3.08) (−2.28) (−1.21) (5.64)

$$(1 - .82205B^{12})^{-1}(1 - B)^2M2_t$$

(15.80)

$$+ (1 - .23012B^2 - .26404B^6 - .23256B^{14} - .15066B^{20})(1 - .59807B)^{-1}(1 - .14054B^{12})^{-1}e_t$$

(4.35) (4.84) (4.18) (2.70) (11.78) (22.43)

RSS = 17.321 RSE = .23266

Residual Diagnostic Check $MQ(24) = 28.184$ (no significant spikes remaining)

$MSL(37) = 50.60$ $MSR(36) = 83.188$ $MSB(73) = 133.79$

No significant spikes were observed between cross correlations of residual and lags of prewhitened M2 except a 0.116 at lag 12 which could not be removed by alternative model specifications that on balance were worse than the present specification. Two significant spikes were observed for cross correlations between prewhitened ($M2/P$) and lags of the residual at 2 and 4 of −0.141 and −0.142, respectively.

Lag	Model Implied Impulse Response Weights	Cumulative Effect
0	−3.416	−3.416
1	−2.148	−5.564
2	−1.350	−6.914
3	3.909	−3.005
4	7.161	4.155
5	4.508	8.663
6	2.829	11.492
7	1.784	13.276
8	1.115	14.392
9	−2.899	11.492
10	−1.819	9.673
11	−1.150	8.522
12	−3.522	5.001
13	−2.219	2.782
14	−1.397	1.385
15	3.040	4.425
16	9.485	13.910
17	5.963	19.873
18	3.745	23.618
19	2.359	25.978
20	1.479	27.457
21	−2.031	25.416
22	−1.280	24.146
23	0.470	24.616
24	−2.007	22.609

Note: For data sources see text. *t*-statistics listed under coefficients. $MQ(24)$ is the modified *Q*-statistic (Ljung and Box, 1976) for the first 24 autocorrelations. $MSL(37)$ is the modified *S*-statistic (Haugh, 1976, p. 383) for 37 cross correlations on the left hand side. The cumulative effect is the sum of the impulse response weights. For further discussion of the method, see text.

rate (which is, however, very close to significance), all other independence tests allow us to reject the hypothesis of independence even with a test that is biased against such a finding. While these results may appear to conflict with the work of Pierce (1977), who reports that he was unable to reject the null hypothesis of independence between demand deposits and the treasury bill rate and between time deposits and the

treasury bill rate, our results are consistent with Pierce's findings that demand deposits are not independent of the federal funds rate and that currency is not independent of the treasury bill rate. As noted earlier, our two filter results for feedback and causality should be interpreted with caution because of a possible downward bias in the modified *S*-statistic, but the significant results obtained are consistent with our more

specific tests of the hypothesis using the one filter procedure.

Results of estimation of the transfer function models are given in table 5. We have calculated the impulse response weights (the $V(B)$ series in equation (11)) and the cumulative effect (the sum of the $V(B)$ weights) and, in the case of the M2 results, have scaled the $V(B)$ vector to make it comparable with the weights obtained for real M2.³⁰

In contrast to the findings obtained using the cross-correlation procedure, the $V(B)$ weights obtained using transfer functions indicate smoother transitions between effects, transitions that are similar to those obtained by Gibson and others using regression analysis. Our results for nominal M2 suggest that after an increase in the money supply the interest rate declines for three periods. Note that the 16 month lag in the positive relationship between money and interest rates found with the one filter method is also found with the transfer function approach. Likewise the transfer function results for real M2 support the somewhat longer lags found in the cross-correlation results for real M2. Furthermore, the lag 7 spike that we rejected using our diagnostic tests for the real money stock does not appear in the transfer function model, indicating that use of our diagnostic procedure enabled us to detect correctly a spurious result.³¹

The modified S -statistic was applied to the residual and lags of the prewhitened input (MSL) and to the prewhitened input and lags of the residual (MSR). Since MSR is significant (52.12 and 83.188) and there are spikes on the right hand side, we conclude that the feedback found with the cross correlation procedure is not spurious. MSL for real balances (39.36) is not significant, indicating that the transfer function model for these series is appropriate.³²

³⁰ The ratio of the means of M2 to M2/P is 117.386. Hence all the $V(B)$ weights calculated for the transfer function for M2 have been multiplied by the scalar 117.386.

³¹ Since the one filter method of cross correlation is one of the ways that a transfer function model can be identified, it is not surprising that these results are similar. What we wish to emphasize is that with the proposed diagnostic test it is possible to detect and, therefore, to disregard spurious cross correlations.

³² MSL for the M2 results remains marginally significant because of a seasonal (lag 12) spike which could not be removed by alternative model specifications.

III. Conclusion

We have presented a method of diagnostic testing for one filter cross correlations. This method has been related to results obtained using the two filter cross correlation independence testing procedure and the transfer function approach. These three methods have been applied in an examination of the relationship between nominal M2 and the commercial paper interest rate and that between real M2 and that interest rate. The results obtained using these three methods are consistent. An advantage of using the one filter method with appropriate diagnostic testing is that liquidity and income effects can be captured separately.

REFERENCES

- Auerbach, Robert, and Jack Rutner, "The Misspecification of a Nonseasonal Cycle as a Seasonal by the X-11 Seasonal Adjustment Program," this REVIEW 60 (Nov. 1978), 601-603.
- Box, C. E. P., and G. M. Jenkins, *Time Series Analysis: Forecasting and Control*, revised edition (San Francisco: Holden-Day, 1976).
- Box, C. E. P., and Larry Haugh, "Identification of Dynamic Regression (Distributed Lag) Models Connecting Two Time Series," *Journal of the American Statistical Association* 72, no. 357 (Mar. 1977), 121-130.
- Brownlee, K. A., *Statistical Theory and Methodology in Science and Engineering*, second edition (New York: John Wiley and Son, 1965).
- Cagan, Philip, *Determinants and Effects of Changes in the Stock of Money 1875-1960* (New York: Columbia University Press, 1965).
- , *The Channels of Monetary Effects on Interest Rates* (New York: Columbia University Press, 1972).
- Culbertson, J. M., "Friedman on the Lag in Effect of Monetary Policy," *Journal of Political Economy* 68, no. 6 (Dec. 1960), 617-621.
- Feige, Edgar, and Douglas K. Pearce, "Economically Rational Expectations: Are Innovations in the Rate of Inflation Independent of Innovations in Measures of Monetary and Fiscal Policy," *Journal of Political Economy* 84 (June 1976), 499-522.
- , "The Casual Causal Relationship between Money and Income: Some Caveats for Time Series Analysis," this REVIEW 61 (Nov. 1979).
- Fisher, Irving, *Appreciation and Interest* (Cambridge: American Economic Association, 1896).
- , *The Rate of Interest* (New York: Macmillan, 1907).
- , *The Theory of Interest* (New York: Macmillan, 1930).
- Friedman, Milton, "The Demand for Money: Some Theoretical and Empirical Results," *The Journal of Political Economy* 67 (Aug. 1959); also reprinted as chapter 6 in *The Optimum Quantity of Money* by Milton Friedman (Chicago: Aldine, 1969).
- , "Interest Rates and the Demand for Money," *Journal of Law and Economics* 9 (Oct. 1966); also reprinted as chapter 7 in *The Optimum Quantity of Money* by Milton Friedman (Chicago: Aldine, 1969).
- , "The Lag in Effect of Monetary Policy," *Journal of*

- Political Economy* 69 (Oct. 1961), 447-466; also reprinted as a chapter in *The Optimum Quantity of Money* by Milton Friedman (Chicago: Aldine, 1969).
- Friedman, Milton, and Anna Schwartz, *A Monetary History of the United States* (Princeton: Princeton University Press, 1963).
- , "Money and Interest Rates," Earlier chapter of forthcoming book *Monetary Trends in the United States and the United Kingdom* by Friedman and Schwartz, University of Chicago memo 13, January 1976.
- , *Monetary Statistics of the United States* (New York: Columbia University Press, 1970).
- Gibson, William, "Interest Rates and Inflationary Expectations: New Evidence," *American Economic Review* 62 (Dec. 1972), 854-865.
- , "Price-Expectations Effects on Interest Rates," *Journal of Finance* 25 (Mar. 1970a) 19-34.
- , "Interest Rates and Monetary Policy," *Journal of Political Economy* 78 (May/June 1970b), 431-455.
- Gibson, William, and George Kaufman, "The Sensitivity of Interest Rates to Changes in Money and Income," *Journal of Political Economy* 76 (May/June 1968), 472-478.
- Granger, C. W. J., "Investigating Causal Relationships by Econometric Methods and Cross Spectral Methods," *Econometrica* 37 (May 1969).
- , "Causality, Model Building, and Control: Some Comments," presented at the IFAC/IFORS International Conference on Dynamic Modelling and Control, July 9-12 (University of Warwick, Coventry), 1973.
- , "Comment on Pierce," *Journal of the American Statistical Association* 72, 357 (Mar. 1977), 22-23.
- Granger, C. W. J., and P. Newbold, "Spurious Regressions in Economics," *Journal of Econometrics* 2 (1974), 111-120.
- , *Forecasting Economic Time Series* (New York: Academic Press, 1977).
- Haugh, Larry D., "Checking the Independence of Two Covariance-Stationary Time-Series: A Univariate Residual Cross Correlation Approach," *Journal of the American Statistical Association* 71 (June 1976), 378-385.
- , "The Identification of Time Series Interrelationships with Special Reference to Dynamic Regression," PhD dissertation, Department of Statistics, University of Wisconsin, Madison, Wisconsin, 1972.
- Johnson, Harry G., "Money in a Neo-Classical One-Sector Growth Model," in *Essays in Monetary Economics* (Cambridge: Harvard University Press, 1967).
- Kuznets, Simon, *Capital in the American Economy: Its Formation and Financing* (New York: National Bureau of Economic Research, 1961).
- Ljung, G. M., and G. E. P. Box, "A Modification of the Overall χ^2 Test for Lack of Fit in Time Series Models," Technical Report No. 477, Department of Statistics, University of Wisconsin, Oct. 1976.
- Mundell, Robert, "Inflation and Real Interest," *Journal of Political Economy* 71 (June 1963), 280-283.
- Nelson, Charles, *Applied Time Series Analysis for Managerial Forecasting* (San Francisco: Holden-Day, 1973).
- Pierce, David, "Residual Correlations and Diagnostic Checking in Dynamic-Disturbance Time Series Models," *Journal of the American Statistical Association* 67 (1972), 636-640.
- , "Relationships—and the Lack Thereof—between Economic Time Series, with Special Reference to Money and Interest Rates," *Journal of the American Statistical Association* 72, no. 357 (Mar. 1977), 11-21.
- Pierce, David, and Larry Haugh, "Causality in Temporal Systems: Characterizations and a Survey," *Journal of Econometrics* 5 (1977), 265-293.
- Sims, Christopher, "Money, Income and Causality," *American Economic Review* 62 (Sept. 1972), 540-552.
- , "Comment on Pierce," *Journal of the American Statistical Association* 72, no. 357 (Mar. 1977), 23-24.
- Tobin, James, "Liquidity Preference and Monetary Policy," this REVIEW 29 (May 1947) 124-131.
- , "The Interest Elasticity of Transactions Demand for Cash," this REVIEW 38 (Aug. 1956), 241-247.