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THE TRANSMISSION OF DATA NOISE INTO POLICY NOISE IN MONETARY CONTROL

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1. INTRODUCTION

In Maravall-Pierce (1983) revisions in the money supply series were analyzed and an attempt was made to measure the effect that revision errors could have on short-run monetary policy. This was done by estimating how often the preliminary measure of the rate of growth of the money supply M1 may give a wrong signal of whether M1 is growing as desired or not, the desired growth being the one lying inside the tolerance range set by the Federal Open Market Committee (FOMC) at each monthly meeting. Using actual data, we computed the number of times a preliminary figure was misleading over the period of the seventies. That frequency turned out to be surprisingly high (close to 40%), and most of the wrong signals could be attributed to seasonal revisions. In fact, the tolerance range used in policy could be interpreted as a relatively narrow confidence interval, under the hypothesis that the final rate of growth of M1 is equal to the preliminary measure. Next, we estimated the probability of a wrong signal under the "ideal" situation in which there are no errors other than seasonal revisions and these revisions are associated with optimal and concurrent seasonal adjustment. This probability was about 20%. Thus, although the proportion of wrong signals could be considerably decreased through improved seasonal adjustment methods, the existence of seasonal revision error sets a non-trivial lower bound to the precision of short run monetary policy.

However, it does not follow that in terms of setting policy the FOMC is necessarily misled by errors in preliminary data. In this paper it is found that noise in the data induces relatively little "noise" in actual policy. The results suggest that the incoming figures are not taken entirely at face value, but rather that in effect a signal-plus-noise separation is made. In fact, we conclude that, on average, for a unit unexpected deviation in the rate of growth of M1 with respect to its target, measured with preliminary data, to a close approximation one-third of that deviation will represent transitory noise which should be ignored, one-third an undesired deviation which should be compensated, associated mainly with money supply shocks, and one-third an unexpected deviation which should be accomodated, primarily associated with shocks stemming from the demand side. It is seen how the different reaction towards demand and supply shocks, together with the signal extraction, explain why noise in the data have little effect on the setting of monetary targets.

The analysis includes some econometric results on errors-in-variables models, which are presented in an Appendix. It should be noted that since 1979, some modifications in monthly operating procedures and in the definitions of the series have been made. As a consequence, the targeted series and the targeting procedures are not identical today to those during the period we consider. However, the present study is still of current interest. Experience with the redefined aggregates is limited and financial innovation is proceeding apace, so that it will be several more years before enough sufficiently well-behaved preliminary data and revisions are available to enable a comparable study to be performed using contemporary aggregates. Of greater significance, the historical statistical characteristics (sizes, variances and autocovariances) of the old and new definitions of the M1 series are broadly similar, the seasonal factor revision process is essentially unchanged, and tolerance ranges akin to those described continue to be used in monetary policy design. Thus, to a considerable degree inference from the 70's experience to the current outlook is warranted.

2. THE DATA SERIES AND THE TARGETS

Since the early 1970's, short run monetary policy has been characterized by the monthly setting of targets for the rate of growth of M1 (seasonally adjusted) over a two-month period, and for the level of the federal funds rate that should prevail until the next FOMC meeting.

The monthly value of M1 seasonally adjusted will be denoted M_t. Let m_t represent the monthly series of annualized rates of growth of M1, seasonally adjusted, calculated for the two-month period beginning with the month t of the FOMC meeting. The series of monthly averages of the Federal funds rate will be denoted r_t.

While the FOMC sets a range of tolerance for m_t without specifying a point estimate, the midpoint of this tolerance range, which we denote $\bar{m}_t$, can be reasonably interpreted as a point target. This interpretation is implied by the wording of the FOMC Record of Policy Action.[1] It is also supported, as we shall see, by empirical evidence. For the funds rate a point target is specified, together with a relatively narrow range, and only occasionally does the point target differ from the midpoint of the range.

To summarize, approximately midway through month t there is a meeting of the FOMC, at which a target $\bar{m}_t$ is set for the growth of M_1 over the months t and (t+1). Also a target for the funds rate is set, which we shall denote $\bar{r}_t$. Thus, in terms of monthly series, when $\bar{m}_t$ and $\bar{r}_t$ are set, information is available up to (and including) month (t-1).

In addition to this short-run target, during most of this period a tolerance range for the long run or annual growth of M1 was usually given as well. The long run targets were set from quarter to quarter, though they

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[1] A typical statement reads "If it appears that the growth rate over the two-month period will deviate significantly from the midpoint of the indicated range, the operational objective for the Federal funds rate shall be modified in an orderly fashion..."

were typically maintained constant for periods longer than three months. We shall use as point target the midpoint of this range, which we shall denote m_t^{LR}. For the first months (when long run targets were not made explicit) the series was set equal to the first available target value.

Since the subscript t refers to the time of the meeting, the first measure of m_t will be available at the meeting held at (t+2). This preliminary estimate, denoted m_t^o, will be revised over a period of approximately three years. When all revisions have been completed, the estimate becomes final and shall be denoted m_t^f. Because of the time needed to complete the revision process, our analysis covers the five year period 1974-78.[2/] The three series $\bar{m}_t$, $\bar{r}_t$, and m_t^{LR} are shown in Figure 1. The series m_t^o and m_t^f are displayed in Figure 2, together with the tolerance range. Finally, Figure 3 shows r_t and its tolerance range.

[2/] The original series and the computation of m_t^o, m_t^f and the intermediate series are described in Maravall and Pierce (1983).

FIGURE 1.

THE SERIES OF TARGETS

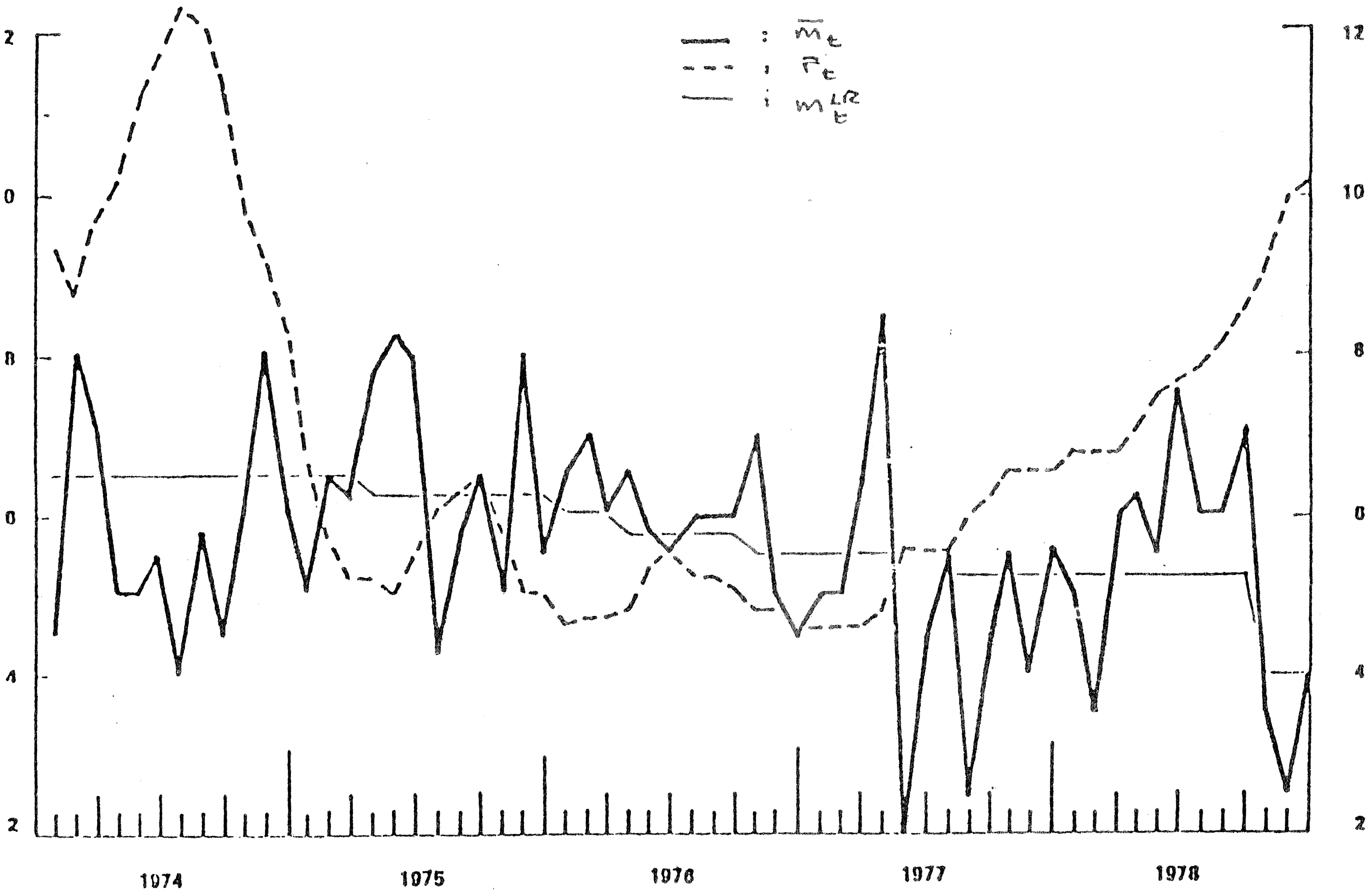


FIGURE 2.

PRELIMINARY AND FINAL MEASURES OF THE RATE OF GROWTH OF M_1

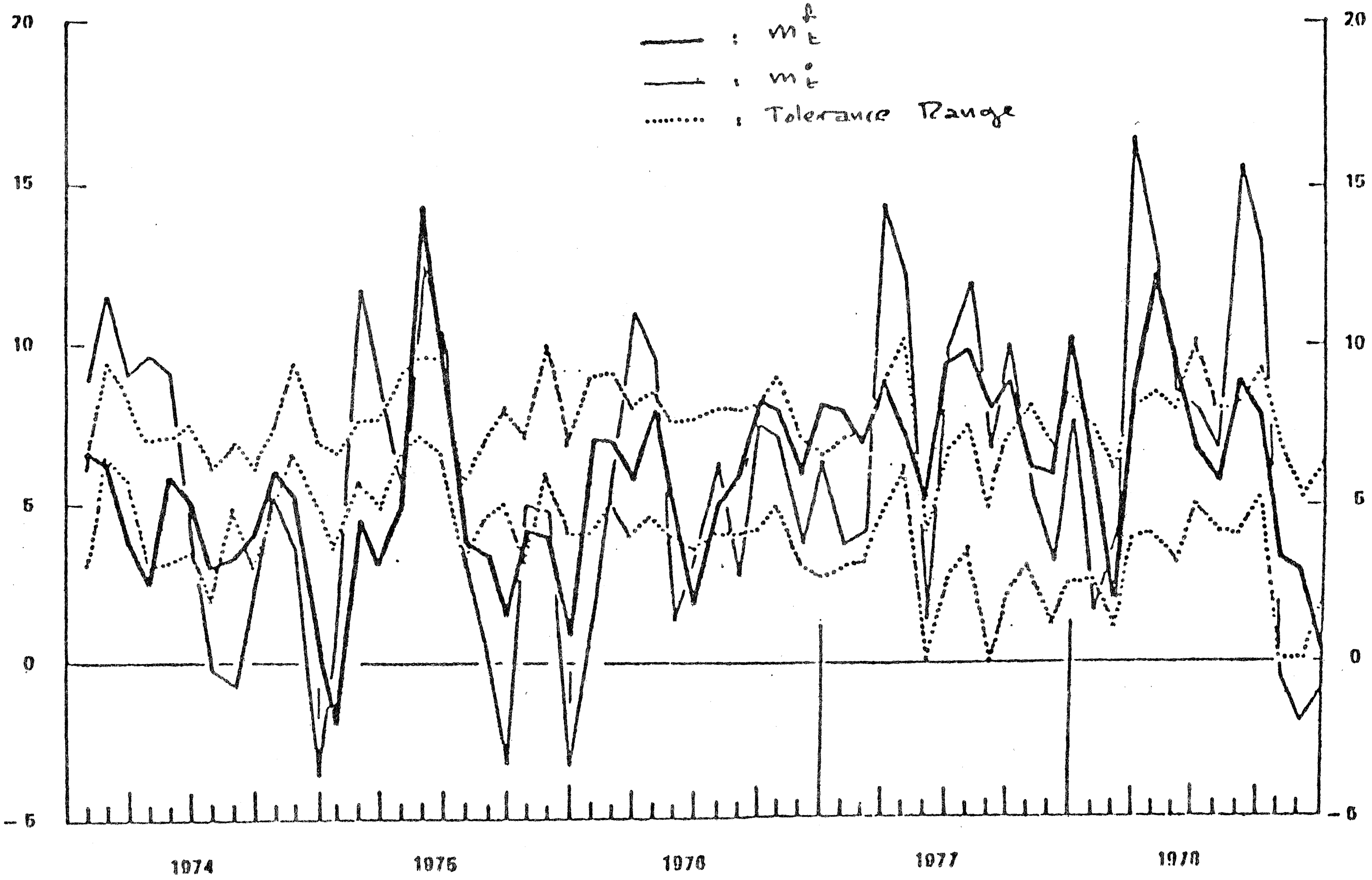
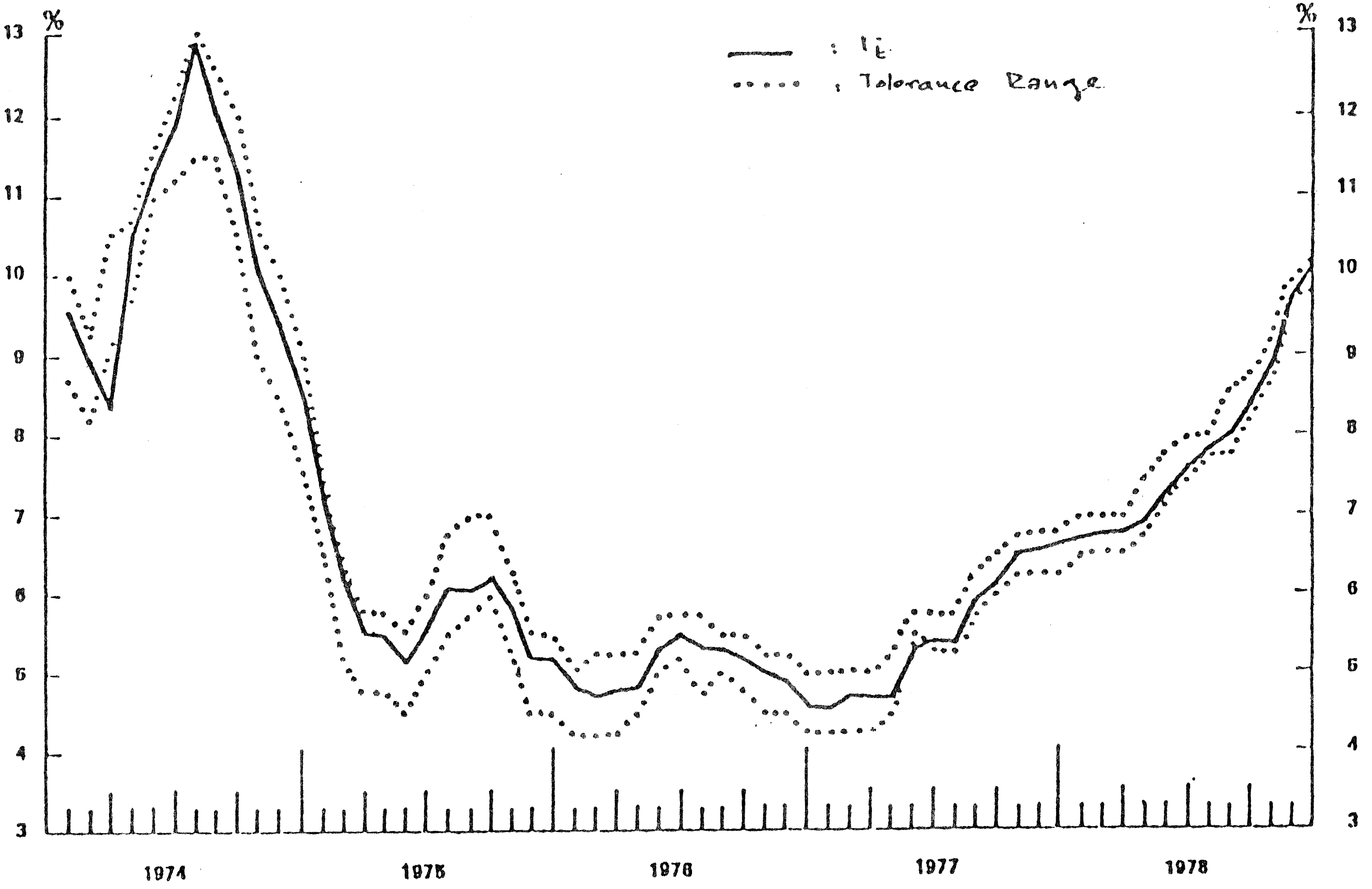


FIGURE 3.

THE SERIES OF FEDERAL FUNDS RATE



3. A MODEL FOR THE SHORT RUN TARGETS

The long run target m_t^{LR} is primarily set in accordance with what is believed to be consistent with such macroeconomic targets as GNP growth, employment, and inflation, and is fairly constant within a year. Thus for now we shall assume that, when setting $\bar{m}_t$ and $\bar{r}_t$, m_t^{LR} is exogenously given. Our model simply states that, when setting short run targets, the FOMC should aim towards the long run one, correcting for undesired deviations as they occur. We shall assume that the full correction extends over a period of several months. Such a gradual response is in agreement with the wording of the FOMC Record of Policy Action. Likely, it reflects mistrust of the preliminary measure on one hand and FOMC concern with orderly markets on the other. This concern typically translates into avoiding unexpected short run fluctuations in the Federal funds rate. (See De Rosa-Stern (1977) and Lombra-Torto (1975).) In this sense, short run targeting should react both to recent deviations in the growth of M1 with respect to its target (so as to be able to meet the long run target) and also to deviations in the funds rate with respect to its target (in order to avoid disorderly markets). Letting

$$dm_t = m_t - \bar{m}_t \qquad (1a)$$

$$dr_t = r_t - \bar{r}_t \qquad (1b)$$

represent both deviations, we shall assume that the targeted money growth rate and interest rate are given, respectively, by

$$\bar{m}_t = \omega(L)\ dm_{t-1} + \lambda(L)\ dr_{t-1} + \gamma m_t^{LR} + u_t \quad . \qquad (2a)$$

and

$$\bar{r}_t = \alpha(L)\ dm_{t-1} + \beta(L)\ dr_{t-1} + \pi m_t^{LR} + v_t \quad . \qquad (2b)$$

where $\omega(L)$, $\lambda(L)$, $\alpha(L)$, and $\beta(L)$ are polynomial distributed lags (DLs) in the lag operator L.

3.1 Supply and Demand Shocks

To a significant extent the targets' variability is assumed to be explainable by whether M1 growth is as desired. We may distinguish three conceptually different reasons for the existence of the discrepancies dm_t between actual and targeted values:

a) an unexpected shock in the money demand function, D_t;

b) an unexpected shock in the money supply function, S_t;

c) an unexpected shock in the IS function.

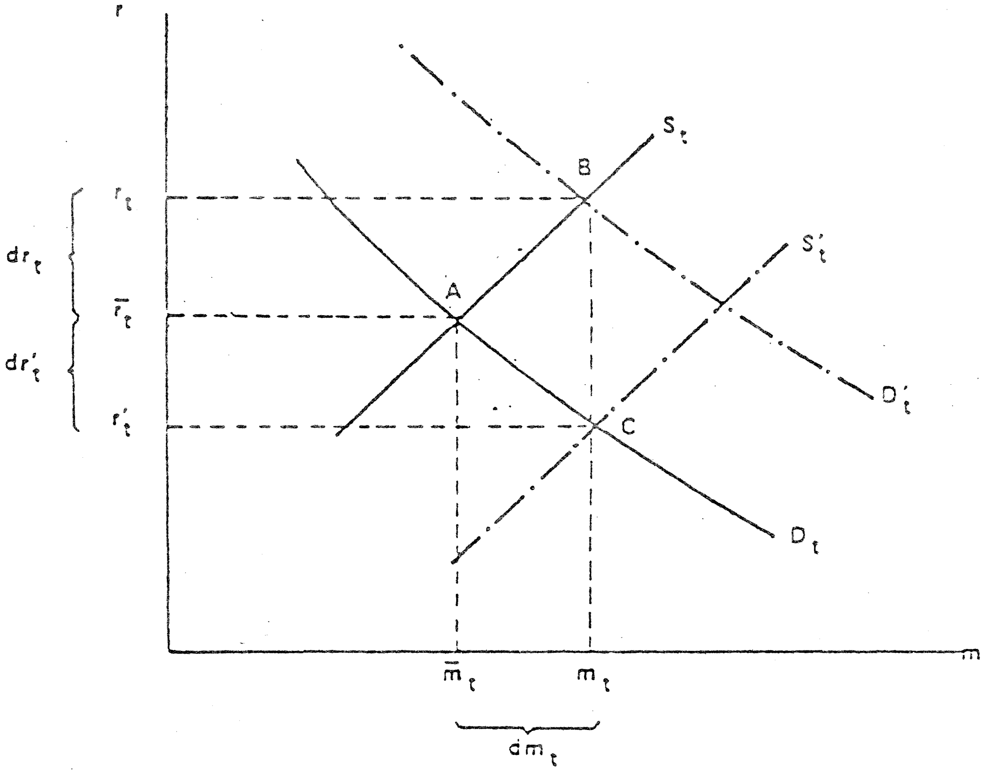
Policy responses to these different shocks should, in principle, be different. (See, for example, Friedman (1977), Davis (1981), and Lindsey (1980), and Section 9 of this paper.) For our purposes, we may group (a) and (c) together, and refer to them as "Demand shocks."

The differential effects of supply and demand shocks are illustrated in Figure 4. The targets set are $\bar{m}_t$ and $\bar{r}_t$, the equilibrium values associated with the demand and supply functions D_t and S_t. Assume there is an unexpected shock in demand, so that D_t moves to D_t'. The equilibrium point attained will then be B instead of A. Hence the deviation in money growth is dm_t and that in interest rates is dr_t. Alternatively, assume the shock affects supply, and S_t moves to S'. The deviations are now dm_t and dr_t'.

If only dm_t is included in equation (2), the two different shifts in Figure 4 would lead to the same response. However, this will not be the case if dr_t is included: an unexpected increase in demand induces a positive dr_t, while an unexpected increase in supply induces a negative dr_t. Hence the two types of shifts can be differentiated.

FIGURE 4.

DEMAND AND SUPPLY SHOCKS



Other variables are of course also relevant in explaining the setting of targets. For example, if monthly FOMC forecasts of income were available, together with monthly measurements, unexpected deviations in income could then be incorporated into equation (2). In principle this would allow us to identify the three different sources of shocks (i.e., to separate money demand from IS shocks). However, some of those relevant variables (such as income) are not observed monthly. Even when monthly values are available, or when monthly estimates can be determined, monthly FOMC targets or forecasts are not available. Reasonably, the informational value of new data to the FOMC depends on the underlying target of forecast implicit in its behavior (Duesenberry, 1977). Hence for analyzing policy such monthly information is of little value.

Equations (2a, 2b) attempt to capture the dynamic reaction to deviations with respect to short run targets. They are in part implied by declared FOMC behavior, and allow for different reactions to supply and demand shocks. Also, through $\bar{r}_t$, $\bar{m}_t$, and m_t^{LR}, policy changes due to shifts in other variables (had they happened or been anticipated) would be incorporated.

Obviously, we cannot expect the equations to account for all variation in targets. But our objective is to estimate how preliminary-data error (noise in data) translates into errors in the setting of the targets $\bar{m}_t$ and $\bar{r}_t$ ("policy noise"). This we hope to capture through the distributed lags on dm_t.

4. THE MODEL IN TERMS OF PRELIMINARY DATA

Obviously, final estimates would be used in the setting of targets if this were possible. Hence equations (2a,2b) should be expressed in terms of final data. However, as described in Section 2, at time t the final values m_t^f and d_{mt}^f $(=m_t^f-\bar{m}_t)$ are unknown, as are the lagged values m_{t-j}^f and dm_{t-j}^f for all j up through the length 12 of the DLs. (We assume that the preliminary observation on r_t has no error). Instead, at time t only a preliminary growth-rate estimate, say m_t°, is available, and the targets must be set using preliminary data.

4.1 Revision Errors

The relation between preliminary and final data may be expressed as

$$m_t^f = m_t^\circ + \delta_t \quad , \qquad (5)$$

where δ_t is the revision error, or the error in preliminary data, due to seasonal and nonseasonal sources, which is corrected in subsequent revisions in the series. It is assumed that δ_t and the preliminary value m_t° are independent, and that δ_t can be expressed as a moving average of future innovations of m_t (see Pierce, 1980).

In addition to equation (5), which relates original and final data via the total revision error δ_t, there are relations involving intermediate revisions of the data. In general, denote by m_{t-k}^k the best estimate of m_{t-k}^f available at time t (which implies, among other things, that concurrent seasonal adjustment is employed). Then the "lag-k revision error," $\delta_{t-k}^{(k)}$ is defined by the relation

$$m_{t-k}^f = m_{t-k}^{(k)} + \delta_{t-k}^{(k)} \quad .$$

Subtracting $\bar{m}_{t-k}$ from both sides in (2), it is then seen that all dm^f_{t-j} appearing as regressors can be written as:

$$dm^f_{t-k} = dm^{(k)}_{t-k} + \delta^{(k)}_{t-k} \quad ,$$

where $dm^{(k)}_{t-k}$ is a function of innovations up to time t, while $\delta^{(k)}_{t-k}$, the revision error still unremoved from the estimate of m_{t-k} at time t, only depends on future innovations. Dropping the superscript k when the context is clear, equation (2a) can be rewritten as:

$$\bar{m}_t = \omega(L)\ (dm^\circ_{t-1}+\delta_{t-1}) + \lambda(L)\ dr_{t-1} + \gamma m^{LR}_t + u_t$$

$$= \omega(L)\ dm^\circ_{t-1} + \lambda(L)\ dr_{t-1} + \gamma m^{LR} + u^*_t \quad , \qquad (6)$$

where

$$u^*_t = u_t + \omega(L)\ \delta_{t-1}$$

is independent of all regressors;3/ similarly for the equation (2b). Thus, in general, (6) would estimate consistently the parameters in (2).

4.2 Benchmark and Seasonal Revisions

At time t, when the FOMC sets $\bar{m}_t$, growth of M_1 for month (t-k) is known for $k \geq 1$; hence m_{t-k}, $k \geq 2$, is known. Hence m°_{t-1} is not known, although it can

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3/ The orthogonality of u^*_t and the regressors in (6) is based on the assumption of concurrent adjustment. However, seasonal revisions are computed once a year. This will introduce some inconsistency in the parameter estimates, but the effect is likely to be small (see Pierce, 1980, and section 9.)

be easily forecasted since the first month of the two-month period is already known. The series of the one-period ahead ARIMA forecasts will be denoted $d\hat{m}^{o}_{t-1}$. (Notice that at time t, $\widehat{dm}^{f}_{t-1} = \widehat{dm}^{o}_{t-1}$.) Again, since

$$dm^{o}_{t-1} = d\hat{m}^{o}_{t-1} + \varepsilon_t$$

where ε_t is orthogonal to $d\hat{m}^{o}_{t-1}$, the use of $d\hat{m}^{o}_{t-1}$ instead of dm^{o}_{t-1} will not pose any serious estimation problem. Also, at time t the FOMC knows dr_{t-k}, $k \geq 1$.

Concerning the revisions in dm^{o}_{t}, it is assumed that

(1) The seasonal revision is "up to date," that is, concurrent adjustment is employed. Because of the once-a-year adjustment used in practice, the value of dm_{t-k} as used in the regression contains a component that will have been revised once. However, this would affect most the more distant regressors, which are likely to be the least important ones.

(2) The non-seasonal revision is removed from the data with an average four-month delay. In this regard, we note that benchmark revisions are made every three months with some additional months of processing involved.[4]

Thus, letting $\tilde{m}^{o}_{t}$ be the rate of growth of the preliminary data corrected for non-seasonal revisions, it follows that the actual regressors in equations (2a,2b) are contained in the vector $[d\tilde{m}^{o}_{t-1}]$ defined by

$$[\widetilde{dm}^{o}_{t-1}] = (d\hat{m}^{o}_{t-1}, dm^{o}_{t-2}, \ldots, dm^{o}_{t-5}, \widetilde{dm}^{o}_{t-6}, \ldots, \widetilde{dm}^{o}_{t-12}) \quad , \qquad (7)$$

where for any t,

$$\widetilde{dm}^{o}_{t} = \tilde{m}^{o}_{t} - \bar{m}_t \quad .$$

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[4] Reasonable alternative hypothesis concerning the delay had practically no effect on the results.

4.3. Final Model

Two further (minor) modifications shall be made to model (2). First, while $\bar{m}_t$ behaves as white-noise, dm_t° and dr_t have low-order autocorrelation and m_t^{LR} is trend dominated. Thus we allow for a lagged endogenous variable. Second, since 12 consecutive monthly releases of M_1 span 13 lagged values of dm_t, an additional term is added to the two DLS. Thus, the final model can be written as the system of two stochastic difference equations

$$(1-\phi_1 L)\bar{m}_t = \omega(L)d\tilde{m}_{t-1}^{\circ} + \lambda(L)dr_{t-1} + \gamma m_t^{LR} + a_t \qquad (8a)$$

$$(1-\phi_2 L)\nabla\bar{r}_t = \alpha(L)d\tilde{m}_{t-1}^{\circ} + \beta(L)dr_{t-1} + \pi m_t^{LR} + b_t \qquad (8b)$$

where the first difference $\nabla\bar{r}_t$ is introduced to remove nonstationarity, $\omega(L)$, $\lambda(L)$, $\alpha(L)$, and $\beta(L)$ are the corresponding DL, and $(a_t,b_t) \sim \underset{\sim}{N}ID(0,\underset{\sim}{\Omega})$ with

$$\Omega = \begin{pmatrix} \sigma_a^2 & \sigma_{ab} \\ \sigma_{ab} & \sigma_b^2 \end{pmatrix}$$

being the contemporaneous covariance matrix. The model (8) has the appearance of a reaction function associated with a loss which depends on the deviations from both the money target and the funds rate target. This pair of equations can also be seen as a reasonable starting model within a COMFAC approach (see Harvey (1981), Chapter 8). Notice that, since $\bar{m}_t$ is a rate of growth, both targets are used in differenced forms, one in logs and one in levels; which is sensible a priori.

The system (8) can be rewritten in the form

$$\bar{m}_t = \omega^*(L)\, d\tilde{m}^\circ_{t-1} + \lambda^*(L)\, dr_{t-1} + \gamma^* \quad m^{LR}_t + u_t \quad , \qquad (9a)$$

$$\nabla\bar{r}_t = \alpha^*(L)\, d\tilde{m}^\circ_{t-1} + \beta^*(L)\, dr_{t-1} + \pi^* \quad m^{LR}_t + v_t \quad , \qquad (9b)$$

where the asterisk denotes the modified DL. Since m^{LR}_t is fairly constant, γ^* and π^* can be assumed to be a constant. The residuals in (9) then follow the AR(1) process

$$\begin{pmatrix} 1-\phi_1 L & 0 \\ 0 & 1-\phi_1 L \end{pmatrix} \begin{pmatrix} u_t \\ v_t \end{pmatrix} = \begin{pmatrix} a_t \\ b_t \end{pmatrix}$$

Equations (8) and (9) are two alternative representations of the model we employ in this paper. The latter has the advantage of directly yielding the distributed lag effects of the exogenous variables; hence the gains (Section 6) are given by $\omega^*(1)$, $\lambda^*(1)$, $\alpha^*(1)$, and $\beta^*(1)$. It also removes from the explained variance that part which is attributable to residual autocorrelation.

There are some constraints that should be satisfied by this model on à priori grounds. Considering (9a), let $m^{LR}_t = m^*(\forall t)$ be an equilibrium constant rate of growth. When $d\tilde{m}^\circ_t = dr_t = u_t = 0$ for all t, consistency of the short and long run targets implies $m_t = m^*$; hence $\gamma^* = 1$. Next, let $v_t = 0$. If, in the absence of external shocks, a constant level of m^*_t implies a constant level for r^*_t, then $\nabla\bar{r}_t = 0$; hence $\pi^* = 0$.

There are additional constraints related to the values of the four gain functions, but we shall discuss them later. The model will be estimated with no constraint and these will then be used as checks on the reasonableness of the results.

5. EMPIRICAL RESULTS

In this section some statistical characteristics of the series of data and targets are established, following which the model of section (4) is estimated.

5.1 Dynamic Structure of the Series

The Autocorrelation Functions (ACFs) of the individual series are displayed in Figure 5. The means and variances appear in Table 1, where the numbers in parentheses represent the asymptotic t-values corrected for the presence of autocorrelation. From the figure it is seen that the money target, $\bar{m}_t$, has the over-all characteristics of univariate white noise. The funds rate target, $\bar{r}_t$, is highly non-stationary, with a strong trend. The variable $\nabla\bar{r}_t$ appears stationary, with some low-order autocorrelation.

The series $[dm_t^\circ]$ measures the difference between actual and targeted growth of M_1 for a two-month period. The target $\bar{m}_t$ is set with information up to (t-1) and refers to growth over periods t and (t+1). It can be interpreted as a two-period-ahead forecast; otherwise policy would be expected to miss the target by some predetermined amount. Specifically, since at time (t+2) the best estimate available of m_t is m_t°, $\bar{m}_t$ can be seen as a two-period-ahead forecast of m_t°. However, it will not be a univariate ARIMA forecast since, when setting targets, the set of information considered by the FOMC is much wider than simply the past values of m_t°. In fact, the ACF of m_t° (see Maravall-Pierce [1983]) resembles that of an MA(1) process. Thus the two-period-ahead ARIMA forecast of m_t° should be close to zero. It follows that dm_t°, given by

$$dm_t^\circ = m_t^\circ - \bar{m}_t \quad ,$$

AUTOCORRELATION FUNCTIONS

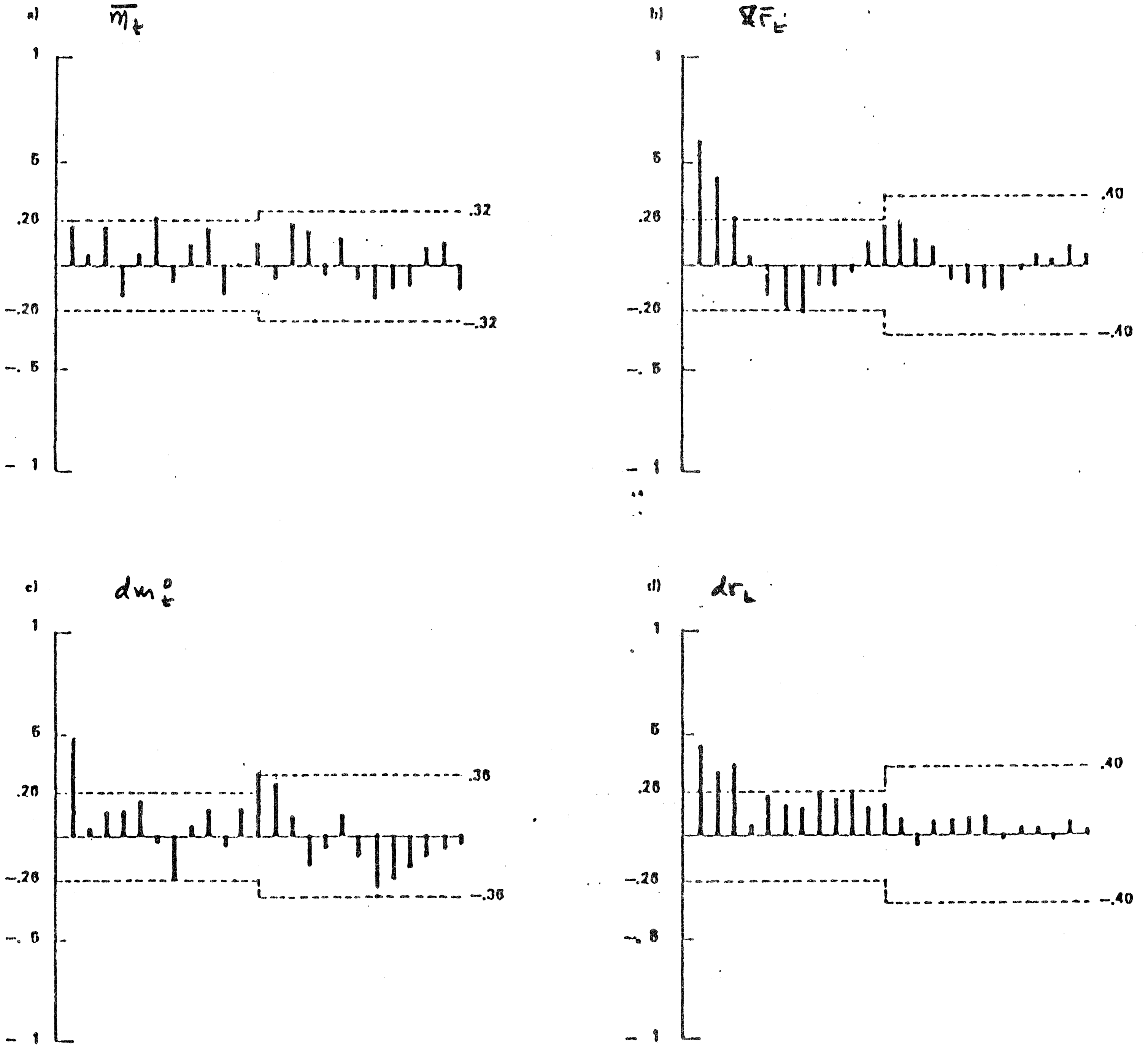


Table 1

Mean and Variance of Target and Target-Deviation Series

Series	$\bar{m}_t$	$\nabla\bar{r}_t$	dm°_t	dr_t
Mean	5.65 (21.7)	.01 (.08)	.24 (.29)	.05 (.98)
Variance	2.02	.28	20.40	.05

should resemble a two-period-ahead forecast error, and hence its ACF should also resemble that of an MA(1) process. This is in agreement with the ACF of the actual $[dm_t^\circ]$ series, except for a small lag-12 autocorrelation.[5/]

The series $[dr_t]$, a monthly series of deviations in monthly averages, can be interpreted, in a similar way, as a series of single-period forecast errors. However, although clearly stationary, the series displays low-order autocorrelation. Evidently there is reluctance in incorporating a systematic component in funds rate misses. Nevertheless, comparing Figures 2 and 3, control of the funds rate appears to have been tighter than that of m_t, although the stardard deviation of dr_t (about 25 basis points) is by no means negligible.

Finally the cross correlation functions (CCFs) between the variables $\bar{m}_t$ and $\nabla\bar{r}_t$, on one hand, and the variables dm_t° and dr_t, on the other, were clearly one-sided, as shown in Table 2 which gives values of

$$\chi^2(10) = T \sum_{k=1}^{10} \rho_k^2 \quad ,$$

ρ_k denoting the lag-k sample cross correlation. This result is in agreement with the hypothesis of exogeneity of dm_t° and dr_t with respect to the targets.

The CCFs were also in agreement with the assumptions made concerning the timing of information (in particular, ρ_0 was close to zero in all cases).

5.2 Estimation of the Model

The model (8) is seen to be in the form of a SURE system, with a common set of regressors, so that OLS is a suitable procedure. Table 3

[5/] Any lag 12 autocorrelation can evidently be attributed to the fact that, 12 periods later, there is a better estimate available, the first-year revision having been performed.

Table 2

χ^2 - Values for Series Cross Correlations

		dm°_{t+k}	dr_{t+k}
$\bar{m}_t$	$k<0$	35.4	35.7
	$k>0$	14.5	4.4
$\nabla\bar{r}_t$	$k<0$	40.8	45.2
	$k>0$	9.6	24.4

summarizes the estimation results. Comparing the variances of the two series of targets (2.02 and 0.28) with the variances of the two series of residuals (0.35 and 0.02), the two equations illustrate a case in which a regression model improves substantially over a univariate time series model for the same series. (Recall that the ACF of $\bar{m}_t$ was close to that of white noise.) The ACF's of the residuals a_t and b_t indicate that both series are essentially white noise, with the corresponding Q-statistics insignificant. Lagrange multiplier tests for the four DL components were carried out, and we detail the derivation for the first of them. Let

$$H_o : \omega(L) \equiv 0 \qquad\qquad (\omega_i = 0, \forall i) \quad ,$$

it is easily seen that, under H_o,

$$a^\circ_t = \bar{m}_t - \phi_1 \bar{m}_{t-1} - \lambda(L) dr_{t-1} - \gamma m^{LR}_t \quad ,$$

$$\left.\frac{\partial a_t}{\partial \phi}\right|_{H_o} = -\bar{m}_{t-1} \quad , \qquad \left.\frac{\partial a_t}{\partial \omega_i}\right|_{H_o} = -d\tilde{m}^\circ_{t-i} \quad ,$$

$$\left.\frac{\partial a_t}{\partial \lambda_i}\right|_{H_o} = -dr_{t-1} \quad , \text{ and } \quad \left.\frac{\partial a_t}{\partial \gamma}\right|_{H_o} = -m^{LR}_t \quad .$$

Then, in the regression of a°_t on $\bar{m}_{t-1}$, $[dm^\circ_{t-1}]$, $[dr_{t-1}]$ and m^{LR}_t ,

$$TR^2 \sim \chi^2_{12} \quad .$$

Thus, it is seen in table 3 that deviations in money growth with respect to its target are highly significant (at the 1 percent level) in the $\bar{m}_t$ - equation, while deviations in the funds rate are borderline at the 10 percent level. For the $\bar{r}_t$-equation, both DL components are highly significant.

Table 3

Summary of Model Estimation Results

		Money Target Equation	Funds Rate Target Equation
R^2		.84	.86
F-statistic		4.51	5.30
Variance of Residuals		$.35 = (.59)^2$	$.020 = (.14)^2$
ACF of Residuals	Q(12)	8.3	11.6
	Q(24)	14.2	17.7
Lagrange Multiplier Test	χ^2_m	28.5	35.9
	χ^2_r	18.3	25.5
Gain of DL	For $d\tilde{m}^{\circ}_{t-1}$	$\omega^*(1) = -.30$	$\alpha^*(1) = .22$
	For dr_{t-1}	$\lambda^*(1) = 1.65$	$\beta^*(1) = -.83$
Coefficient of m_t^{LR}		$\gamma = .73$ (4.24)	$\pi = 0$ (.63)
Coefficient of m_{t-1}		$\phi_1 = .28$ (1.72)	$\phi_2 = .27$ (1.41)

(t-values are given in parenthesis.)

5.3 Role of the Federal Funds Rate

The difference in the significance of funds-rate deviations in the two equations may have a reasonable explanation. Assume, for example, that at time (t-1), $\bar{m}_{t-1}$ and $\bar{r}_{t-1}$ are set and that, being in equilibrium, $\bar{m}_{t-1} = m_t^{LR}$. Further assume that, shortly after the meeting, incoming data indicate that m_{t-1} will be larger than desired. If by increasing the funds rate (within the tolerance range) the growth of M_1 is brought back to the desired path, then there would be no reason to modify $\bar{m}_t$ at the next meeting, assuming m_t^{LR} remains unchanged. What could be expected is $\bar{m}_t = m_t^{LR}$ and $\bar{r}_t > \bar{r}_{t-1}$. In this case, m_t would not depend on dr_{t-1}, while obviously r_t would.

On the other hand, if the increase in the funds rate needed to bring m_t to the desired path is judged too large, then some deviation in money growth would be accepted and, likely, $\bar{m}_t < \bar{m}_{t-1}$ (in order to meet m_t^{LR}) and $\bar{r}_t > \bar{r}_{t-1}$. Thus, although $\bar{m}_t$ may depend on dr_{t-1}, this dependence is stronger for the case of $\bar{r}_t$. If, eventually, the target is not met and $dm_{t-1} \neq 0$, then both targets would be modified. Hence deviations in money growth should be significant in both equations.

5.4 Estimated Lag Distribution

The shapes of the four DLs are given in Figure 6. The ω^*-weight tends to decrease as the corresponding lag increases, except for a small peak at lag 12.[6] The λ^*-weights behave following a more erratic pattern, in accordance with the fact that $\lambda(L)$ was estimated with less precision. The

---

[6] This peak could be attributed to the fact that some of the more distant lagged dm°_{t-1} values would not be used as such when setting targets, since the first year seasonal revision would already be available. Roughly, what is likely to happen is that some corrections are made after 12 months of additional data have become available.

α^* and β^*-weights both gradually decrease, exhibiting negative correlation between adjacent values. This correlation is also present in $\omega(L)$ and $\lambda(L)$ and can be attributed to the lag-1 autocorrelation in the $[dm_t^o]$ and $[dr_t]$ series. The correlation between adjacent coefficient estimates within a particular DL is not a matter of concern to us since we shall not be interested in individual coefficient estimates. Of more interest to us are the values of the gain functions, or total multipliers.

6. THE GAIN FUNCTIONS

Deviations of monetary aggregates from their targets were seen in Section 3 to be caused by different types of unanticipated shocks, for which different policy responses were appropriate. During the period we consider, FOMC intended behavior was, for a supply shock, adherence to the monetary aggregate target, offsetting therefore the money growth deviation (see Lindsey, 1980 and references in section 3.1.). For a shock originating on the demand side, a more adequate response was considered to accomodate, at least partly, the change.

Assume an equilibrium situation, satisfied at all times before t, when there have been no shocks, and $dm^{\circ}_{t'} = dr_{t'} = 0$, $t > t'$. Such a system is growing at the rate

$$\bar{m}_{t'} = m^{LR}_{t'} = m^* \quad ,$$

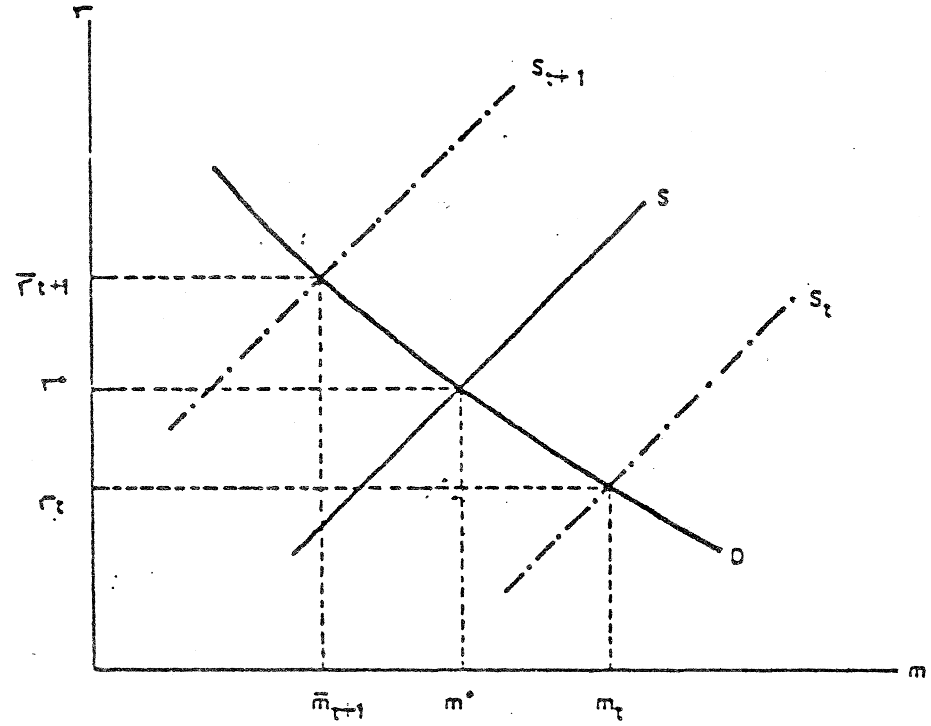
with the funds rate set at the level

$$\bar{r}_{t'} = r_{t'} = r^* \quad ,$$

so that $\nabla\bar{r}_{t'} = 0$. The values (m^*, r^*) are constant and are the equilibrium values associated with an underlying supply-demand system, S and D, such as in Figure 7.

FIGURE 7

RESPONSE TO MONEY SUPPLY SHOCKS



At time t, $\bar{m}_t = m^*$ and $\bar{r}_t = r^*$, but assume there is an unexpected (one-period) shift in supply, so that S moves to S_t. The new equilibrium values will be m_t and r_t, hence

$$dm^\circ_t = m_t - m^* > 0 \quad , \quad dr_t = r_t - r^* > 0 \quad . \qquad (11)$$

Assume that m is computed over periods of one month, and that deviations with respect to targets are also offset in one month. The exogenously given long-run target (measured quarterly) remains unaffected, so that in period (t+1) the monetary authority will have to decrease the money supply in such a way as to compensate for the undesired supply shocks. Thus the authority will attempt to move the supply toward S_{t+1} and the new targets will be $\bar{r}_{t+1}$ and $\bar{m}_{t+1}$. Therefore, $\bar{r}_{t+1} > r^*$ $(=\bar{r}_t)$, and $m_{t+1} < m^*$ $(=m^{LR}_t)$, so that

$$\nabla\bar{r}_{t+1} > 0, \quad \bar{m}'_{t+1} < 0 \quad , \qquad (12)$$

where $\bar{m}'_{t+1} = \bar{m}_{t+1} - m^{LR}_{t+1}$. Moving from the comparative statics framework to our model written as (10), expressions (11) and (12) imply that the total multipliers should satisfy the constraints

$$\omega^*(1) < 0 \quad , \quad \lambda^*(1) > 0$$

and

$$\alpha^*(1) > 0 \quad , \quad \beta^*(1) < 0 \quad .$$

If the unexpected shift is in demand, insofar as it is partly offset, similar reasoning yields

$$\omega^*(1) < 0 \quad , \quad \lambda^*(1) < 0$$

and

$$\alpha^*(1) > 0 \quad , \quad \beta^*(1) > 0 \quad .$$

Thus in both cases the gain of the DL which applies to deviations in money growth has the same sign, while the one corresponding to deviations in the funds rate has different signs according to whether the unexpected shift is in demand or in supply. Therefore:

1. $\omega^*(1)$ should be smaller than zero. However, it is difficult to specify a priori a numerical value, since such a value depends on the relative importance of the deviations that are accomodated. We shall simply require

$$-1 \leq \omega^*(1) \leq 0 \quad . \tag{13a}$$

2. Since the numerical value of $\alpha^*(1)$ depends on the units of measurement of money growth and interest rates, we simply require

$$\alpha^*(1) \geq 0 \quad . \tag{13b}$$

3. The expressions $\lambda^*(1)$ and $\beta^*(1)$ should have opposite signs, although which is positive depends on whether supply shocks or demand shocks dominate in the short run. Thus

$$\mathrm{sgn}\ |\lambda^*(1)| = -\ \mathrm{sgn}\ |\beta^*(1)| \quad . \tag{13c}$$

The four values of the gains are displayed in Table 3, where it is seen that the constraints (13) are satisfied.

Also, from the signs of $\lambda^*(1)$ and $\beta^*(1)$, it is evident that the short run is mostly characterized by supply shocks [7].

---

[7] Poole (1977) states "my guess is that the vast bulk of weekly and monthly money-growth surprises reflect money-supply disturbance rather than either IS or money-demand disturbances".

In fact, it is for this type of disturbances that intermediate money stock targeting is more addecuate (see Davis [1981]).

Given that in the short run money supply is more volatile than money demand, we would expect negative correlations both between the residuals a_t and b_t of (8) and between the targets $\bar{m}_t$ and $\nabla\bar{r}_t$. In fact the estimated cross-correlation between a_t and b_t is $\rho_0 = -.22$ and that between the targets is $\rho_0 = -.30$.

Finally, the coefficients of m_t^{LR} are also in agreement with the "a priori" values. For the m_t equation, $\gamma = .72$, and $\phi_1 = .28$, so that $\gamma^* \simeq 1$. For the $\nabla\bar{r}_t$ equation, $\pi = 0$, hence $\pi^* \simeq 0$.

FIGURE 8

ACTUAL AND FITTED VALUES OF m_t.

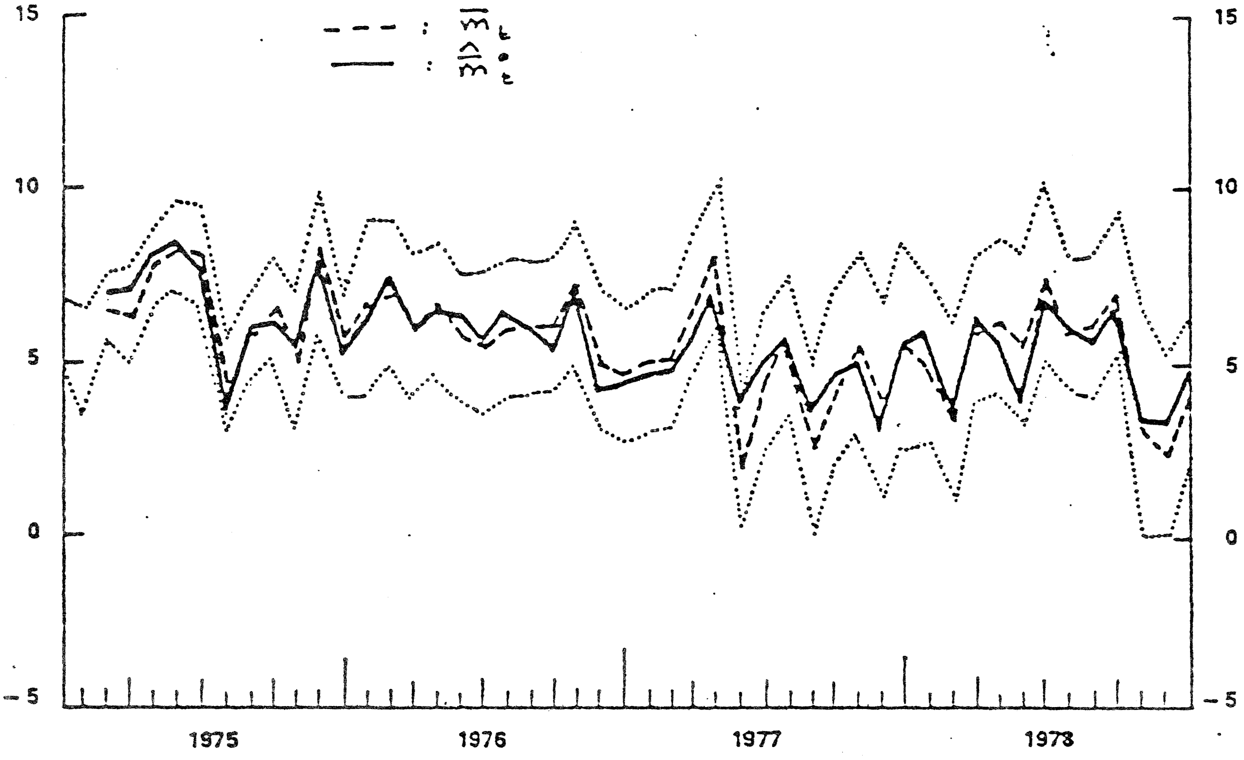
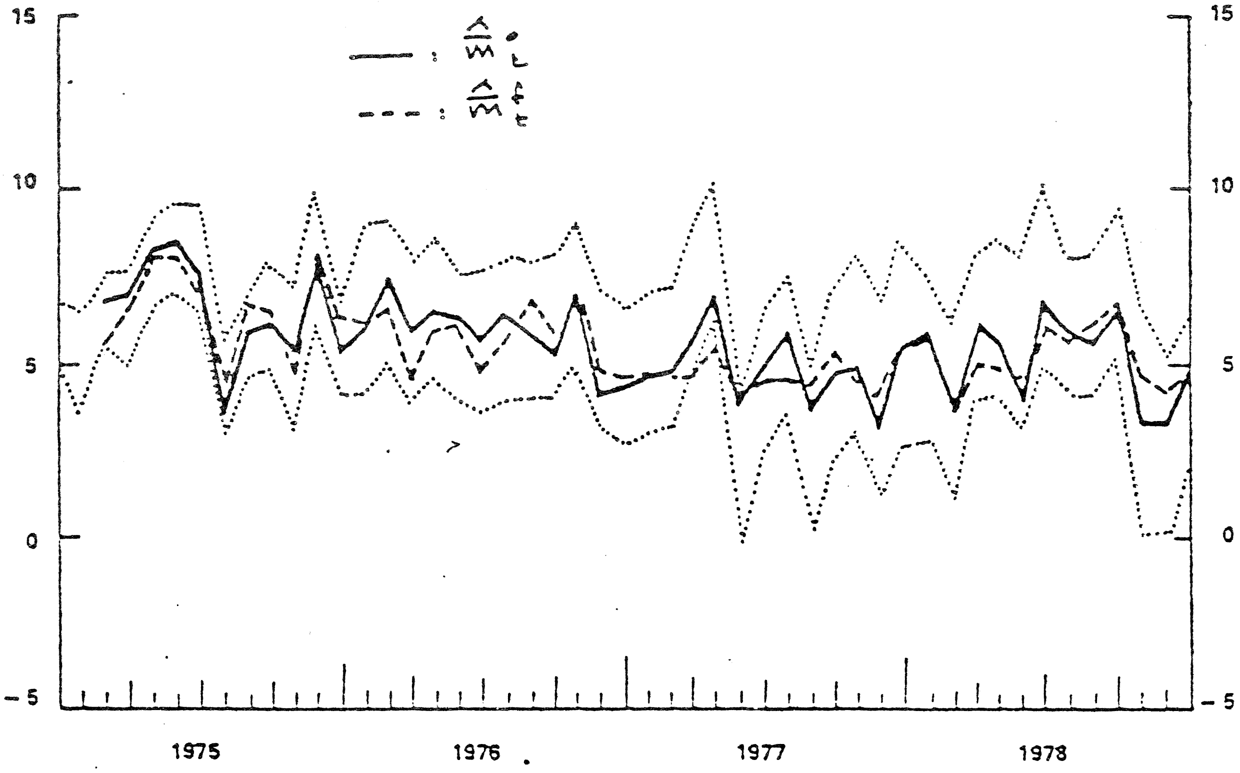


FIGURE ~~9~~ 10

FITTED VALUES WITH PRELIMINARY AND FINAL DATA (series m_t)



7. THE TARGETS WITH PRELIMINARY AND FINAL DATA

Orthogonality of the revision error and the preliminary money stock measure imply that the model estimated with preliminary data can be used as the model that would be applied be fitted to the final data 8/. Thus we can infer what the targets would have been if the final data had been known.

First, for the preliminary data, using the sequence $\{d\tilde{m}^\circ_t\}$ and setting $\hat{\phi}_1 = .30$, $\hat{\phi}_2 = 0$, fitted values of $\bar{m}_t$ and $\bar{r}_t$, denoted $\bar{m}^\circ_t$ and $\bar{r}^\circ_t$, were computed through (9) with $u_t = v_t = 0$.

Second, replacing the sequence $\{dm^\circ_t\}$ in (7) with the corresponding one for final data (the first element being $d\hat{m}^f_{t-1}$, the one-period ahead ARIMA forecast), we obtain the estimate of targets that would have been set if final data had been available, $\bar{m}^f_t$ and $\bar{r}^f_t$. Figures 8 and 9 show the series of actual and fitted targets, together with the (actual) tolerance ranges.

In order to assess the effect of the revisions, we note first that the difference between the two sets of fitted values is

$$x_t = \hat{m}^f_t - \hat{m}^\circ_t = -\omega^*(L)\ \delta_t \tag{15}$$

Similarly, for the funds rate the estimated revision effect is

$$y_t = \hat{r}^f_t - \hat{r}^\circ_t = -\ \alpha^*(L)\delta_t \quad . \tag{16}$$

Note that in (15) and (16), δ_t is δ°_t and $L^j\delta_t$ is $\delta^{(j)}_{t-j}$.

Figures 10 and 11 compare $\{\bar{m}^\circ_t\}$ with $\{\bar{m}^f_t\}$, and $\{\bar{r}^\circ_t\}$ with $\{\bar{r}^f_t\}$, respectively. Practically all the targets that would have been set if final data had been available lie within the tolerance range, set when only preliminary information on recent money grow is available.

8/ In our case, the parameters of the "final" model would asymptotically efficiently estimated with OLS applied to the preliminary data.

FIGURE 8

ACTUAL AND FITTED VALUES OF m_t.

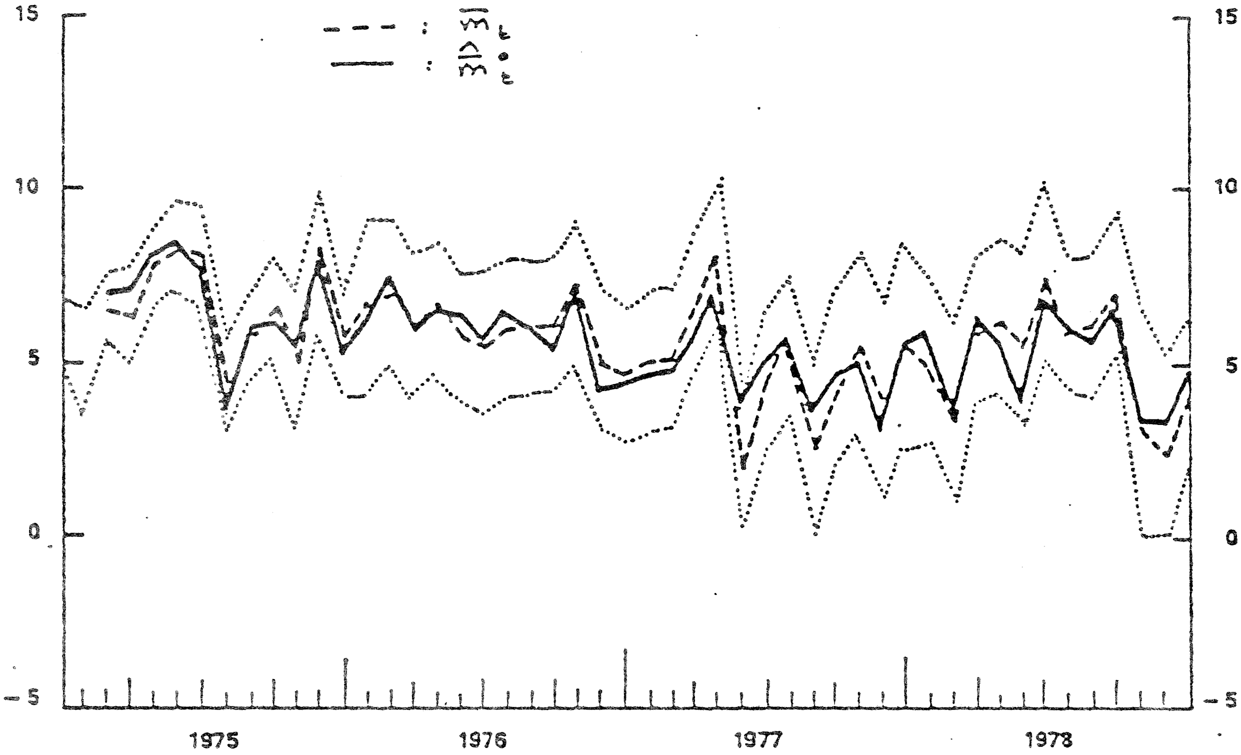


FIGURE 9

FITTED VALUES WITH PRELIMINARY AND FINAL DATA (series m_t)

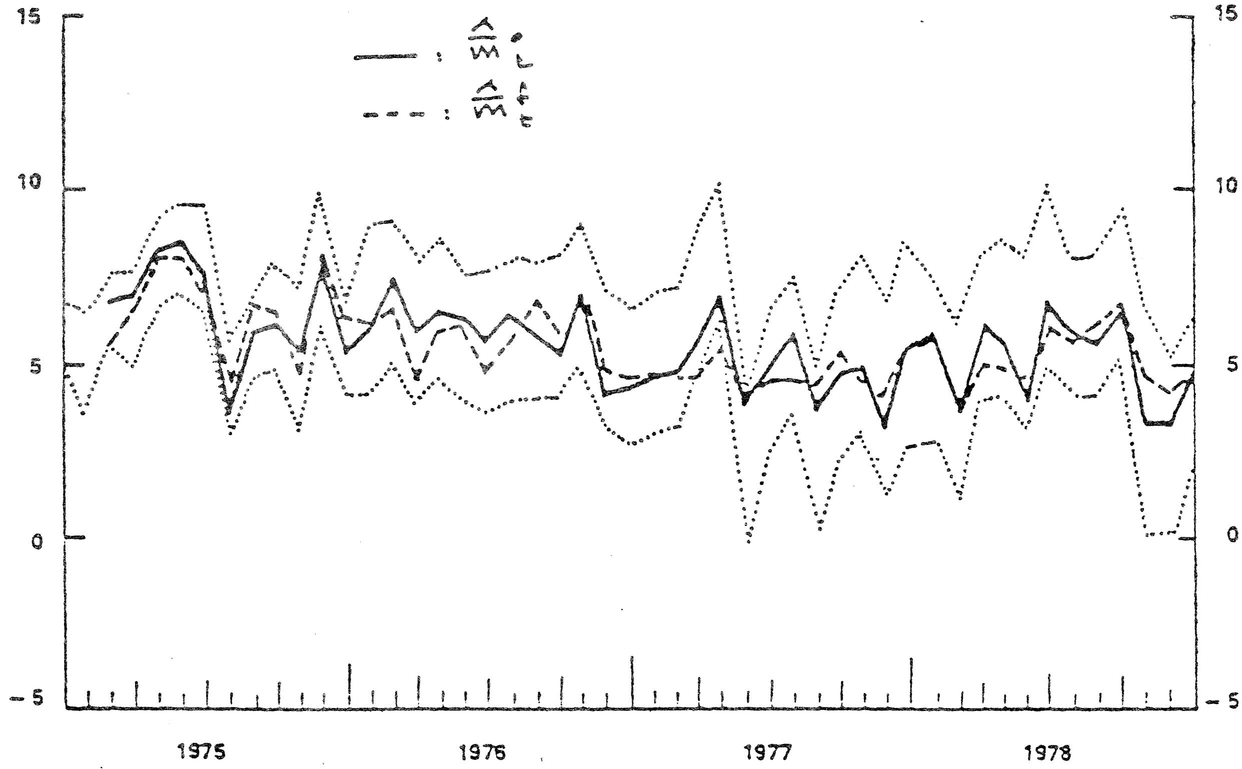


FIGURE. 10

ACTUAL AND FITTED VALUES OF r_t.

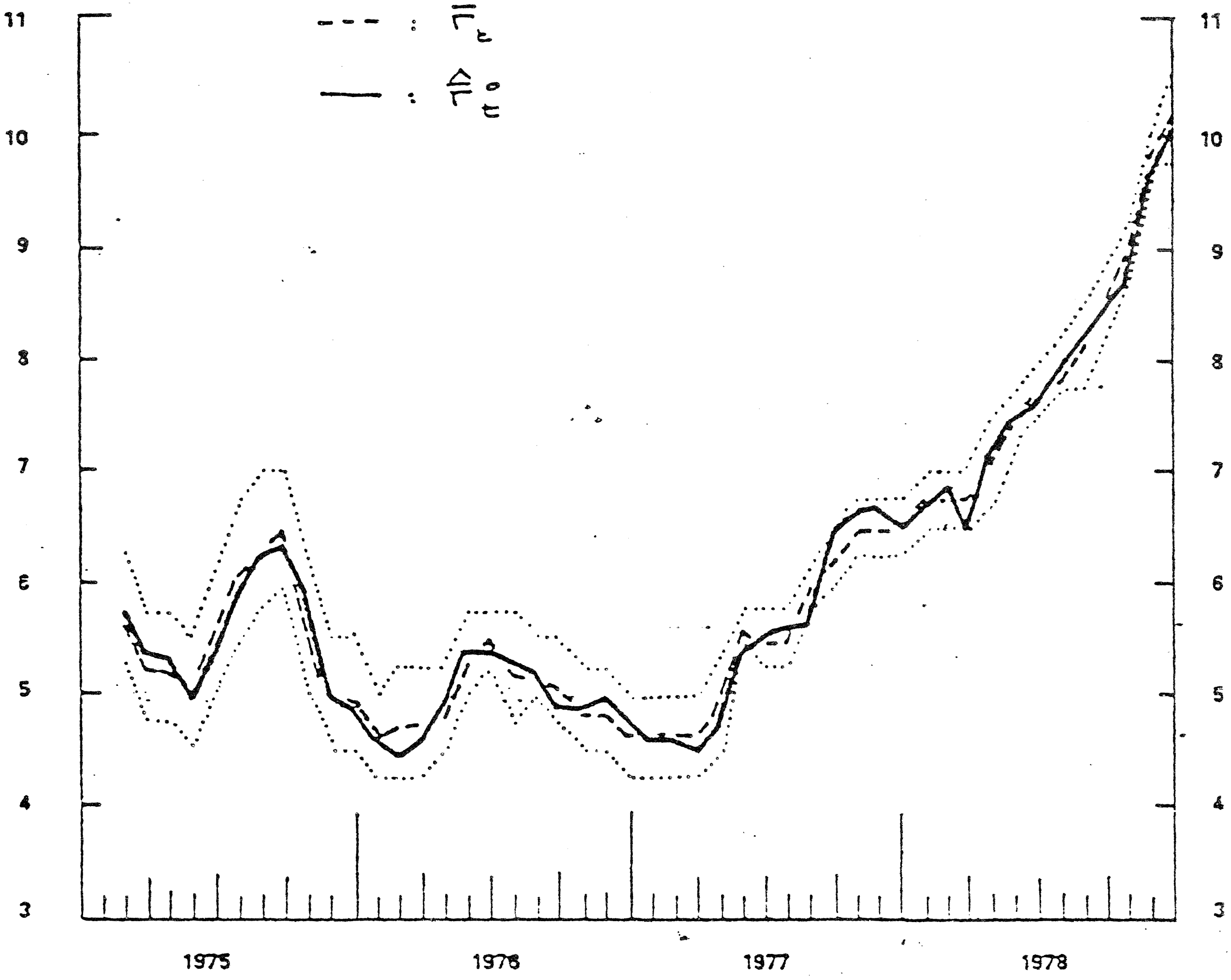
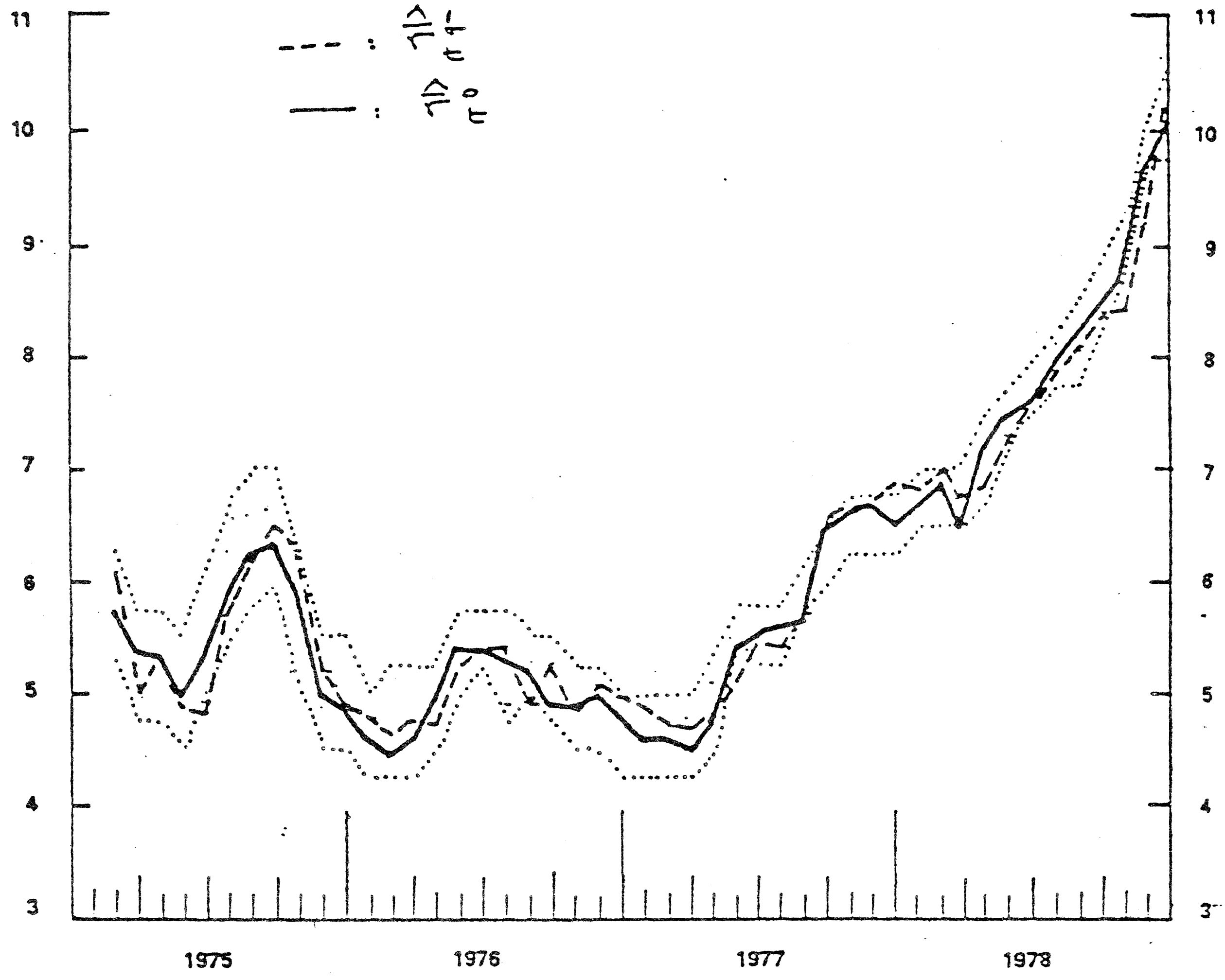


FIGURE - 11

FITTED VALUES WITH PRELIMINARY AND
FINAL DATA (series r_*)



Comparing Figures 2 and 10, it is seen that, although preliminary and final data often give conflicting signals as to whether growth of M1 is as desired or not, the effect of these conflicting signals on the setting of short-run targets is rather small. The targets would not have been much different if the (revision) error in preliminary data would not have been present, in spite of its size.

This smoothing effect is also evidenced by the fact that, while the standard deviations of δ_t and δ^s_t, the total and seasonal revision errors, are 3.57 and 2.72, the standard deviation of the difference in m-targets is .69 .[9] The smoothing, according to (15), is due to $\omega^*(L)$. To get a better understanding of this mechanism, let us assume first that all deviations dm_t are fully offset (a "pure" monetary aggregate targeting policy). If z_t denotes the annualized monthly rate of growth of M1 (in percent points), so that

$$z_t = 1200 \nabla \log M_{t+1} \quad ,$$

then since

$$m_t = 600(1+L)(1-L)\log M_{t+1} \quad ,$$

it follows that

$$m_t = \frac{1+L}{2} z_t \quad .$$

If implicit in the two-month targets there are monthly targets, then

$$dm_t = \frac{1+L}{2} dz_t \quad , \qquad (17)$$

where dz_t represents the monthly deviation. Hence the term "$\omega^*(L)dm_t$" of (10a) becomes:

$$W(L)dz_t = \frac{1}{2}(1+L)\omega^*(L)dz_t \qquad (18)$$

---

[9] Notice that, since δ_t and $\bar{m}_t$ have the same units, $\omega^*(L)$ is a-dimensional. The rest of the discussion concentrates on the money growth target equation.

It is easily seen that, if an undesired change in the level of M1 is offset by exactly the same amount, then

$$W(1) = -1$$

Hence setting L = 1 in (18) yields

$$\omega^*(1) = -1 \quad .$$

However, our estimate of $\omega^*(1)$ was -.30, which implies that, for a deviation of 1, only .30 of it would eventually be offset.[10] It seems quite unlikely that money demand shocks that should be accomodated can account for 70 percent of target misses. An explanation of the low value of $\omega^*(1)$ is given in the next section.

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[10] The standard deviations of the estimates of $\omega^*(1)$ and $\alpha^*(1)$ were .15 and .04, respectively.

8. NOISE IN FINAL DATA

Growth of the money supply is subject to erratic, transitory movements that tend to cancel out over relatively short periods (see Pierce _et al_, 1981). Such movements are present in both demand and supply shocks. In terms of policy, it could be reasonable to ignore them, focusing instead on a smoother component, presumably some type of trend.[11] Thus assume the FOMC intends to react to a signal μ_t^f in the final data, where

$$m_t^f = \mu_t^f + n_t^f \quad ,$$

and the noise n_t^f is orthogonal to the signal. If consequently the targets are set for the signal, then

$$dm_t^f = d\mu_t^f + n_t^f \quad ,$$

where $d\mu_t^f = \zeta_t^f - \bar{m}_t$. In terms of the preliminary data,

$$dm_t^\circ = d\mu_t^\circ + n_t^\circ \quad , \tag{19}$$

where $d\mu_t^\circ$ and n_t° are the undesired deviation of and the noise in the preliminary signal. From (5), if ν_t is the signal component in δ_t ,

$$d\mu_t^f = d\mu_t^\circ + \nu_t \quad ,$$

The noise-reduction effect then follows easily: The revision in the data is large, but the revision in the signal contained in it is relatively small. Since targets are set for the signal, the difference in targets induced by revisions in the data

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[11] A similar argument has been made for other variables also followed closely by policy makers. (See, for example, Blinder [1980] and Davidson [1982].)

is also small.12/ This explanation is also in agreement with the dynamic features of the series of differences in targets, x_t, and of revisions, δ_t. Figure 12 compares the ACF of x_t and δ_t. The shape of both functions is similar; hence the large difference in variance can be attributed to a large noise component in the revision series.

If the preliminary growth measure is not taken at face value and the targets express the desired growth of a signal contained in the data, then in equation (9a) the variable dm°_t should be replaced by $d\bar{\mu}^\circ_t$; and similarly for (9b). Having used dm°_t, we have incurred in a traditional Errors-in-Variables (EIV) situation, and the error u_t is correlated with dm°_t (through n°_t). Therefore, our parameter estimates are inconsistent.13/ However, since our purpose was to compare fitted values obtained with (9a), our interest is in estimating $E_t(\bar{m}_t)$, which is the conditional expectation of $\bar{m}_t$ given $[dm^\circ_{t-1}]$, $[dr_{t-1}]$, and $[m^{LR}_t]$. It is shown in the Appendix that this conditional mean is correctly estimated by using OLS on (9a).14/ In other words, the OLS inconsistent estimates applied to the noisy data provide consistent estimates of $E_t(\bar{m}_t)$. Therefore, the comparison we performed between $\bar{m}^\circ_t$ and $\bar{m}^f_t$ (i.e., the change in targets if the final data had been available) is still valid, despite the EIV structure of the model. As for the estimates of the parameters in $\omega^*(L)$ and of the gain $\omega^*(1)$, it is also shown in the Appendix that the effect of the EIV is to reduce their value by a factor inversely proportional to the signal-to-noise ratio.

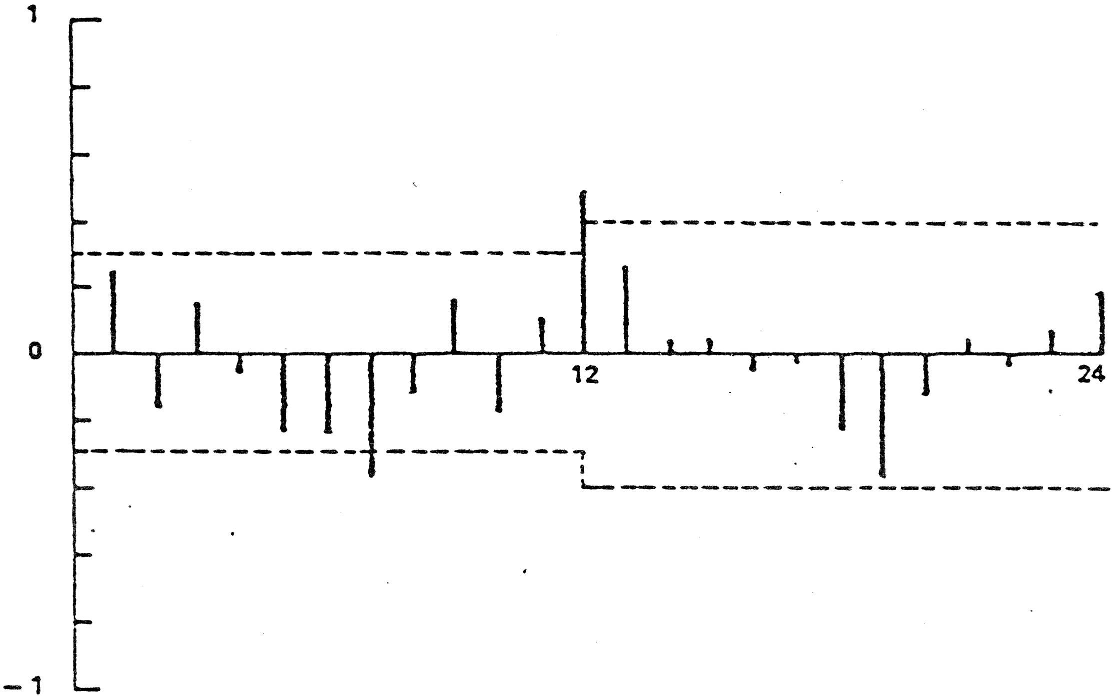
12/ This interpretation is, on occasion, contained in press coverage of monetary policy. For example, the lack of reaction to self-cancelling noise is implied by the following quotation: "The Fed is viewing the April M-1 growth as an aberration, and is willing to give it some time to be reversed in coming weeks" (International Herald Tribune, May 3, 1982). The use of a signal which is less affected by revisions is implicit in the following excerpt from an editorial in The Washington Post: "The rule of wisdom, ... for people who make policy, is to pay more attention to general trends over the months than to the latest flash number. An unexpected number may mean that a trend is changing. Then again, as time passes, it may also be the number that gets changed" (March 17, 1982).

13/ Since the regressors $[d\tilde{m}^\circ_t]$, $[dr_t]$, and $[m^{LR}_t]$ are approximately orthogonal, the estimates of $\lambda^*(L)$ and γ^* will be unaffected by the presence of error in dm°_t (see the Appendix).

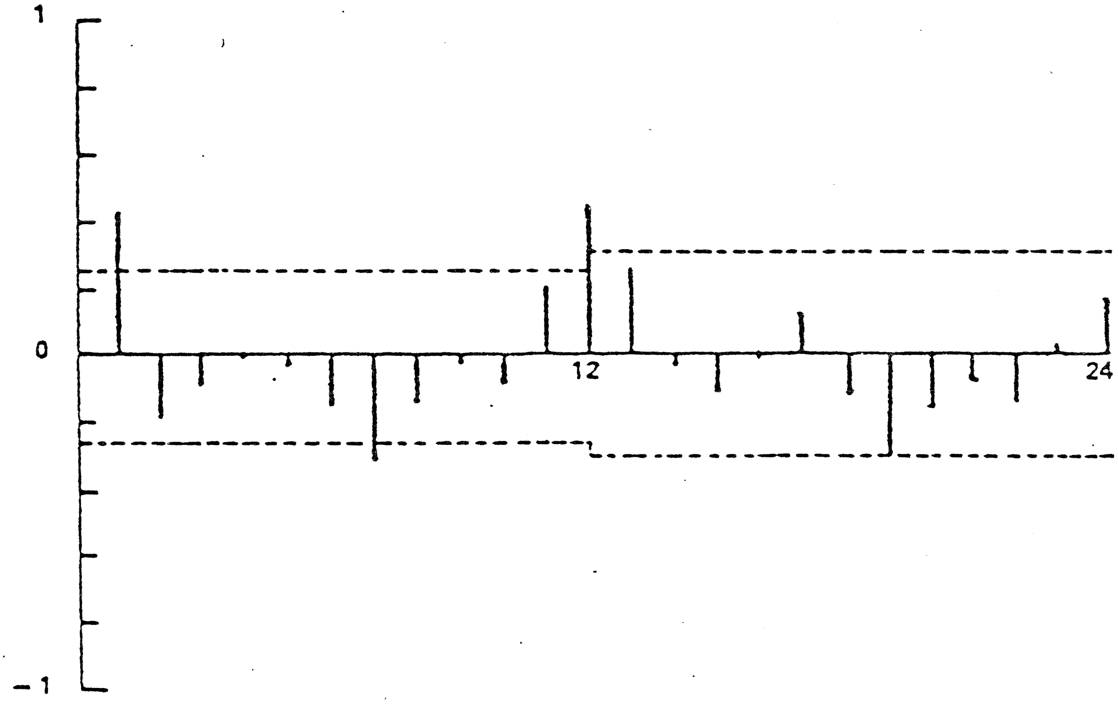
14/ More generally, we show that EIV assumptions do not cause any harm to either the fits or the forecasts.

FIGURE 12

ACF OF $[x_t]$



ACF OF $[d_t]$



An analysis of the noise component of the M1 series, based on univariate statistical techniques, is contained in Pierce et al (1981), where it is assumed that whatever part of the M1 series is serially uncorrelated should be considered transitory noise. While such a method of noise extraction is unlikely to be used in the conduct of monetary policy, where considerations other than past values of M1 are also relevant, it is interesting to compare our results with their findings. For the month-to-month rates of growth of M1 (seasonally adjusted), the estimated standard deviation of the noise is 4.5; when the two-month rates of growth are considered, this estimate becomes 2.5. The resulting variance ratio (see equation A.8) is .31, so that the plim of the estimate of the gain would be attenuated by a factor of (1-.31), or .69.

Consequently, for a unit unexpected deviation in the rate of growth of M1, .30 would be offset and .31 could represent irrelevant transitory movements. The rest then represents deviations due to money demand shocks that should be accomodated, plus biases due to other sources of error, such as that implied by the once-a-year instead of concurrent seasonal adjustment. (Based on results reported in Bayer and Wilcox (1981), a reasonable value for the asymptotic downward bias on the estimate of the gain would be in the order of .1.). Therefore, even if most transitory deviations represent supply shocks, the proportion of demand shocks that are accomodated seems large. Since accomodation of (non-transitory) demand shift would eventually show up in changes in m_t^{LR}, the steady decrease of m_t^{LR} over our period may be partly related to a downward move of money demand associated with institutional and technological changes in financial markets (see Simpson and Porter, 1980).

To summarize, roughly 1/3 of M1 target misses can be attributed to transitory noise, 1/3 to deviations that are offset, and 1/3 to deviations that are accomodated. It is the extraction of noise and the accomodation of the demand-induced deviations that explains why relatively large revisions in the data have relatively small impact on the setting of targets.

9. A FINAL COMMENT ON THE MEASUREMENT OF THE GROWTH OF M1

It has been seen that the effective series on which monetary policy is based can be viewed as the result of a smoothing of a two-month rate of growth of seasonally adjusted M1. Since self-cancelling, transitory noise should be removed irrespective of the source of the shock, it makes sense, first to remove the noise and seasonality, extracting from the series a signal (presumably, some type of trend), and second, to identify which part of the deviation in the signal should be accomodated.

The point may be quite relevant. For example, if the two-month targets are assumed to hold for the first of the two months (all targets expressed as annualized percent points), then the ACF of the series of monthly deviations in preliminary data resembles that of white noise, with variance of 38.85. Using as an estimate of the variance of the noise the one in Pierce (1981), equal to 20.25, the ratio of the signal variance to the series variance is .52. If the series of deviations is white noise and the signal and noise independent, the latter two also have to be white noise. Hence, by a well-known result,

$$E(d\mu_t^\circ | dm_t^\circ) = .5 \, dm_t^\circ$$

where $d\mu_t^\circ$ and dm_t° are as in (16), but for month-to-month deviations. Thus, prior to any policy response, a new preliminary measured monthly deviation should be cut in half.

Finally, while seasonal adjustment relies heavily on statistical estimation, noise extraction is mostly judgemental. However, in general, signal (or trend) extraction within a model based approach offers several advantages. First, it facilitates systematic analysis, hence methodological improvements. Second, it could simplify seasonal adjustment, avoiding possible inconsistencies in the present procedure. Finally, it makes "political bias" more difficult to use. This bias is reflected in a tendency to consider a large undesired increase in M1 a statistical aberration when interest rates are high, and a large decrease an indication of the FOMC commitment to anti-inflation policy in periods of high inflation.

APPENDIX: SOME RESULTS ON EIV MODELS

Let the model be

$$y = X\underline{\beta_1} + Z\underline{\beta_2} + u \quad ,$$

where Z represents a set of variables observed without error and X is not directly observable. Instead, observations are available on a variable W related to X by

$$W = X + V \quad . \tag{A.1}$$

The shock u is assumed NID$(0,\sigma_u^2)$, uncorrelated with X, Z, and V. The errors in V are uncorrelated with X and Z. All random variables are assumed Normal with zero mean. The variables X, V, and Z have finite limiting variance-covariance matrices Σ_X, Σ_V, and Σ_Z, and Ω, Σ, and Σ_W denote the variance-covariance matrix of (W,Z), (X,Z) and W, respectively.

Let $\underline{b} = (\underline{b_1}', \underline{b_2}')'$ be the OLS estimators of $\underline{\beta} = (\underline{\beta_1}', \underline{\beta_2}')'$ in the regression of y on (W,Z). Then it is well-known (see, for example, Levi, 1973) that

$$\text{plim } \underline{b} = \Omega^{-1} \, \Sigma\underline{\beta} \quad .$$

If W and Z are uncorrelated, it is easily seen that

$$\text{plim } \underline{b} = \text{plim} \begin{pmatrix} \underline{b_1} \\ \underline{b_2} \end{pmatrix} = \begin{pmatrix} \Sigma_W^{-1} \; \Sigma_X \; \underline{\beta_1} \\ \underline{\beta_2} \end{pmatrix} \tag{A.2}$$

and hence the OLS estimator of $\underline{\beta_2}$ is consistent. In what follows, we shall not consider variables measured without error.

Consistency of OLS Fits

Let the true model be:

$$y = X\underline{\beta} + u \quad ,$$

where (A.1) holds, together with the relevant assumptions of the previous section. For a particular set of observations W, consider the estimation of $E(y|W)$, and let $\underline{b}$ denote the OLS estimator of $\underline{\beta}$ in the regression of y on W.

Lemma:

$$E(y|W) = W \operatorname{plim} \underline{b} \quad .$$

Proof:

$$\begin{aligned} W \operatorname{plim} \underline{b} &= W \operatorname{plim}(W'W)^{-1}W'(X\underline{\beta} + u) = \\ &= W \operatorname{plim} \left[(W'W)^{-1}W'X\right]\underline{\beta} \quad . \end{aligned}$$

Since (W,X) are jointly Normal,

$$E(X|W) = W \Pi \quad ,$$

where Π is estimated consistently by $(W'W)^{-1}W'X$, (i.e., by an OLS regression of X on W). Hence

$$W \operatorname{plim} \underline{b} = W\pi\underline{\beta} = E(X|W)\underline{\beta} = E(y|W) \quad , \qquad \text{q.e.d.}$$

The result tells us that, although an EIV assumption produces inconsistent OLS parameter estimates, it does not cause much harm to the OLS fits as estimators of the conditional mean of y. The inconsistent parameters estimators applied to the noisy data provide consistent estimators of the expected value of the endogenous variable, for a given set of observations.

(For the case in which X has one variable,

$$E(y|W) = B\,E(x|W) = \beta \frac{\sigma_x^2}{\sigma_w^2} W \quad .$$

Since $$\operatorname{plim} \underline{b} = \frac{\sigma_x^2}{\sigma_w^2} \beta ,$$

$$E(y|W) = \operatorname{plim} \; b \; w \quad .)$$

The proof is easily extended to show that

$$E(y|W_f) = W_f \operatorname{plim} \underline{b} \quad ,$$

where W_f represents out of sample values of the exogenous variables. Thus the Lemma applies equally to forecast computation.

Effect on Individual Parameters and on the Gain

From the first subset of equations in (A.2),

$$\operatorname{plim} \underline{b} = \Sigma_W^{-1} \Sigma_X \underline{\beta} \quad , \qquad \text{(A.3a)}$$

and, since $\Sigma_W = \Sigma_X + \Sigma_V$, this can also be expressed as

$$\operatorname{plim} \underline{b} = (I - \Sigma_W^{-1} \Sigma_V)\underline{\beta} \quad . \qquad \text{(A.3b)}$$

For the model we consider in the paper, W denotes lagged values of the observed variable dm_t°. The true unobservable variable X is the signal $d\mu_t^\circ$, and V is the noise n_t°, where

$$dm_t^\circ = d\mu_t^\circ + n_t^\circ \quad , \qquad \text{(A.4)}$$

with uncorrelated signal and noise. In section 5 we saw that dm_t° could be assumed to be an MA(1) process. In fact, since the lag-1 autocorrelation of

dm_t° is approximately .5, the MA parameter should be equal to 1. This is also in agreement with the following argument. From (14),

$$dm_t^\circ = \frac{1}{2}\,(dz_t + dz_{t-1}) \quad ,$$

and hence

$$dm_t^\circ = \frac{1}{2}\,[(z_t - \bar{z}_t) + (z_{t-1} - \bar{z}_{t-1})$$

Since both are successive 1-period ahead forecast errors, it follows that dm_t° has approximately the MA(1) representation

$$dm_t^\circ = \frac{1}{2}\,(a_t + a_{t-1}) \quad , \qquad (A.5)$$

where a_t is white-noise. Thus dm_t° is a non-invertible MA(1) process.

The three variables in (A.4) are monthly series of two-month periods. Hence each series "overlaps" one month. The noise contained in one month appears in two successive values of n_t°. Thus n_t° should also be an MA(1) process. Since dm_t° is also an MA(1), the same should be true of the signal $d\mu_t^\circ$. Therefore, the three variables in (A.4) are MA(1)'s. The following Lemma allow us to identify uniquely the parameters in the signal and noise process.

Lemma: Let y_t be the sum of several independent components, each an MA(1) process. If y_t is non-invertible, then all components are also non-invertible.

Proof: Write

$$y_t = \sum_i x_t^{(i)} \quad ,$$

where

$$x_t^{(i)} = a_t^{(i)} + \theta_i\, a_{t-1}^{(i)}$$

with $a_t^{(i)} \sim NID(0, \sigma_i^2)$. Obviously y_t is an MA(1) process, of the type

$$y_t = b_t + \theta b_{t-1} \quad .$$

Since y_t is non-invertible, $\theta = 1$ or $\theta = -1$. Consider first the case $\theta = 1$. Then:

$$\rho_y(1) = \frac{\sum_i \theta_i \sigma_i^2}{\sum_i (1+\theta_i^2)\sigma_i^2} = \frac{1}{2} \quad . \tag{A.6}$$

Letting $k_i = \sigma_i^2/\sigma_1^2$, $i > 2$, expression (A.6) yields

$$(1-\theta_1)^2 + \sum_{i \geq 2} (1-\theta_i)^2 k_i = 0 \quad . \tag{A.7}$$

Since $k_i > 0$, $\forall i$, this implies $\theta_i = 1$, $\forall i$.

When $\theta = -1$, the 1/2 of expression (A.6) becomes -1/2 and (A.7) is replaced by

$$(1+\theta_1^2) + \sum_{i \geq 2} (1+\theta_i)^2 k_i = 0 \quad ,$$

which implies $\theta_i = -1$, $\forall i$, finishing the proof.[15]

Applying the Lemma to our model, we have that the three series in (A.4) are MA(1) processes with unit root, and parameter equal to 1. Therefore, expressions (A.3a and b) can be greatly simplified. The three matrices Σ_w, Σ_x, and Σ_y can be expressed as hH, where, in all cases, H is the matrix:

$$H = \begin{pmatrix} 1 & .5 & 0 \ldots 0 \\ .5 & 1 & .5 \cdots 0 \\ 0 & .5 & \ddots \quad 5 \\ \vdots & \ddots & \ddots \;\; \vdots \\ 0 & 0 \ldots .5 & 1 \end{pmatrix}$$

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[15] An obvious corollary is the following: under the same assumptions, if there is at least one invertible component, the aggregate will also be invertible.

and h is σ_w^2, σ_x^2, and σ_v^2, respectively. Thus, after simplification (A.3) becomes

$$\text{plim } \underline{b} = \frac{\sigma_x^2}{\sigma_w^2} \underline{\beta} = \underline{\beta} \left(1 - \frac{\sigma_v^2}{\sigma_w^2}\right)$$

Therefore the EIV assumption has a constant effect on each of the β parameters. The effect is to shrink the numerical value towards zero.

In terms of the gain, letting

$$g_b = \underline{1}'\underline{b}$$

$$g_\beta = \underline{1}'\underline{\beta} \ ,$$

where $\underline{1}' = (1...1)$, it is easily seen that

$$\text{plim } g_b = \frac{\sigma_x^2}{\sigma_w^2} g_\beta = g_\beta \left(1 - \frac{\sigma_v^2}{\sigma_w^2}\right) , \qquad \text{(A.8)}$$

so that the same shrinking effect takes place. The net effect is seen to depend on the relative contributions of the signal and noise to the variance of the observed series. Since

$$\frac{\sigma_x^2}{\sigma_w^2} = \frac{1}{1+\nu} \ ,$$

where $\nu = \sigma_x^2/\sigma_v^2$, the asymptotic bias can be expressed in terms of the signal-to-noise ratio. The smaller the signal, the larger the bias will be.

In the terminology used in the paper, $g_b = -.30$, $g_\beta = \omega^*(1)$, $\sigma_\omega^2 = \text{Var } (dm_t^\circ)$ and $\sigma_v^2 = \text{Var } (n_t^\circ)$.

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