Monetary vs. Credit Aggregates: An Evaluation of Monetary Policy Targets*

W. DOUGLAS McMILLIN
Louisiana State University
Baton Rouge, Louisiana

JAMES S. FACKLER
University of Kentucky
Lexington, Kentucky

I. Introduction

The past decade was characterized by increased emphasis upon monetary aggregates as intermediate targets for monetary policy.[1] Although growth rate ranges were specified for both narrow and broad monetary aggegates, as well as for bank credit, the primary focus of the Federal Reserve was on the growth rate of narrow money. However, some policy makers [19] have recently expressed doubt about the viability of using narrow money as an intermediate target. In particular, it has been argued that recent financial innovations (such as the introduction of the various NOW accounts, the proliferation of money market mutual funds, and the adoption of cash management techniques by both firms and consumers, to name a few) have distorted the *M1/GNP* relationship, thereby raising the question of whether broader monetary aggegates or even credit aggregates might bear a closer relationship to income. Furthermore, the Federal Reserve has recently announced growth rate target ranges for a broad credit aggregate in addition to growth rate ranges for *M1* and *M2*.[2] Friedman [7;8] provides a general discussion of these issues while Modigliani-Papademos [18] provide a theoretical analysis of the conditions under which a credit aggregate is preferred to a monetary aggregate as an intermediate target. However, nothing conclusive can be said at the theoretical level since the choice between money and credit aggregates depends upon the true underlying structure of the macroeconomy. Hence the choice among money and credit aggregates is an empirical issue.

The aim of this paper is to provide evidence on the issue of credit aggregates versus monetary aggregates as intermediate targets for monetary policy. The paper evaluates the characteristics of several monetary and credit aggregates in terms of two attributes of a

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1. For a review of Federal Reserve operating techniques in the postwar period see Wallich-Keir [27].

2. The announcement of target ranges for the broad credit aggregate (defined later in the text) appeared in [26,39].

good intermediate target. These attributes are: (1) a consistent, predictable relationship to the ultimate goals of policy and (2) the absence of significant feedback from non-policy variables to the target variable.

One method of evaluating the closeness and reliability of the relationship between alternative money and credit variables and the ultimate policy goals is the estimation of separate "St. Louis" reduced-form equations in which a proxy for the ultimate policy goals—the growth of nominal *GNP*—is regressed on distributed lags of current and past growth rates of a financial aggregate and a fiscal policy variable like high-employment government expenditures.[3] A comparison of results across the equations containing different financial aggregates could then be made. However, these "reduced-form" models have been the subject of much criticism [3;9;10;17]. On the basis of these criticisms, use of the single equation approach in this study was eliminated as a fruitful way of evaluating the monetary and credit aggregates. However, in order to keep the analysis at a manageable level, we focus on the variables used in the single equation approach (nominal *GNP*, a financial aggregate, and a fiscal policy variable) in a multi-equation analysis.

An alternative to the single equation approach that enables one to evaluate both attributes of the potential intermediate targets is the recently proposed vector autoregressive modeling technique of Hsiao [11;12]. The vector autoregressive (VAR) model—which is described more fully in section II—can be viewed as a system of reduced form equations with a separate equation for each variable in the system. Thus the models evaluated in this paper contain an equation for nominal *GNP*, a financial aggregate, and a fiscal policy variable (high employment expenditures). In specifying the models, no a priori assumptions are made about the exogeneity of the policy variables or the lag length for the variables in each equation. The specification of each model provides evidence on the exogeneity of both the financial aggregate and the fiscal variable[4] and the estimated systems can be employed to analyze the effects upon nominal *GNP* of a change in either the financial aggregate or the fiscal policy variable.

The modeling technique is discussed in section II and the trivariate models are presented in section III. The models are discussed in section IV and the final section contains concluding comments.

II. Estimation of the Money and Credit Models

The methodology used to estimate the trivariate models analyzed here is the vector autoregressive technique suggested by Hsiao [11;12] and extended by Caines, Keng, and Sethi [1]. The VAR modeling technique is employed rather than a single equation or a structural model approach, since the VAR models avoid imposing potentially spurious a priori constraints (such as, for example, exogeneity of the monetary variable in the nominal *GNP* equation) on the model. The VAR technique employed involves the use of the Granger-causality definition in conjunction with Akaike's final prediction error criterion to impose restrictions on the estimation of the VAR. The technique allows each variable to depend

3. For a representative study that estimates "St. Louis" reduced-form equations see Seaks-Allen [25].

4. An additional reason for employing the vector autoregressive approach is based upon the results of Nelson and Schwert [20]. Using Monte Carlo simulations they found that the most powerful causality tests are those based on the correct reduced-form model and that parametric tests involving reduced-form models are more powerful than tests employing cross-correlations of univariate ARMA residuals or regressions based on univariate ARMA residuals.

upon a subset of the variables in the system and allows for different lag lengths for each variable in each equation.

The VARs estimated in the manner described above differ from the unconstrained VARs estimated by Sims [22;23], Fischer [6], and Dwyer [4] in which each variable depends upon all other variables in the system with the same lag length. One problem that emerges in the estimation of a Sims-type VAR is that lengthening the common lag by one increases the number of parameters to be estimated by the square of the number of variables and thus rapidly depletes the degrees of freedom available for estimation. The degrees of freedom problem becomes significant in estimating Sims-type systems since the lag length must be kept generous in order to avoid under-specifying the lag for one or more of the variables and thereby avoiding biased coefficient estimates. Further, there is no reason to believe that the same lag length is appropriate for all variables in each equation.

The use of VARs to analyze the money and credit variables is motivated by Fischer's [6,402] observation that vector autoregressions are ". . . a convenient way of summarizing empirical regularities and perhaps suggesting the predominant channels through which relations work." Furthermore, Sims [24,138], in a discussion of his VAR results, notes that ". . . careful attention to the historical data exerts an important discipline on what can be plausibly asserted about the way the economy works." However, the VAR is a reduced-form technique and it is thus difficult, based upon VAR results, to distinguish sharply among structural hypotheses.

The Lucas critique is also potentially applicable to the VAR technique, and it is assumed in this paper that there were no changes in policy regimes over the estimation period. As Sims [24,138] notes, "The U.S. postwar data contain enough information to give a useful characterization of the conditional distribution of the future of major macro-economic aggegates given the past. Although there is evidence that this structure changes over time, there is also evidence that it does not change suddenly, so that a model fit to the whole postwar period as if its parameters were fixed over that whole period is not badly biased because of parameter changes."

The Hsiao procedure is illustrated by discussing the specification of the nominal *GNP* equation in the models estimated; a more technical discussion is provided in an appendix which is available from the authors on request. Since the theory underlying the estimation of the VAR is based upon the use of stationary data [12;21], the first step in the Hsiao procedure is to suitably transform the data to achieve stationarity. The specific transformations employed in this study are discussed in section III; at this point it is sufficient to emphasize that stationary data are used.

The next step in specifying the nominal GNP equation in a typical model estimated is the determination of the own lag length for GNP. This is done by varying the lag in the autoregression $Y_t = a_0 + a_{11}(L)Y_t + e_t$ from 1 to m where Y = nominal GNP transformed to be stationary, $a_{11}(L)$ is a distributed lag polynomial such that $a_{11}(L) = \Sigma_{k=1}^{m} a_{11k} L^k$, L is the lag operator so that $L^k Y_t = Y_{t-k}$, m = highest order lag (specified a priori to be 15)[5], and e_t = zero mean white-noise error term. The *FPE* is calculated for each autoregression and is defined for lag k, k=1, . . . , m, as

$$FPE_{(k)} = [(T+k+1)/(T-k-1)][SSR_{(k)}/T]$$

5. If the optimal lag turns out to be 15 (or even 13 or 14) then prudence suggests that the procedure should be rerun with the lag extended beyond 15 in order to determine if the optimal lag is longer than 15.

where T = number of observations used in estimating the autoregression, and SSR = sum of squared residuals. The lag length that minimizes the FPE is selected as the order of $a_{11}(L)$.

Hsiao [12] points out that the FPE criterion is equivalent to using an F-test with a varying significance level. As Judge, et al. [13] note, an intuitive reason for using the FPE is that an increase in the lag length increases the first term but decreases the second term and these opposing forces are balanced when their product reaches a minimum. Thus, according to Hsiao [12,88], the FPE criterion is "... appealing because it balances the risk due to the bias when a lower order is selected and the risk due to the increase of variance when a higher order is selected."

Once the order of $a_{11}(L)$ is found, a determination of whether the financial and fiscal variables enter the Y equation is made. The procedure begins with the estimation of the bivariate equation $Y_t = a_0 + a_{11}(L)Y_t + a_{12}(L)X_t + e_t$ where X_t = relevant financial and fiscal variables (considered one at a time), and $a_{12}(L)$ is a distributed lag polynomial defined in a similar manner to $a_{11}(L)$. $a_{11}(L)$ is fixed at its previously determined order (k) and the lags in $a_{12}(L)$ are varied over $l = 1, \ldots, m$. The FPEs for the resulting equations are computed and are defined for lag l, $l = 1, \ldots, m$, as

$$FPE_{(k,l)} = [T+k+l+1)/(T-k-l-1)][SSR_{(k,l)}/T].$$

The lag length for X_t that yields the minimum FPE is selected as the lag order for that variable. This FPE is then compared to the FPE from the previous step. If min $FPE_{(k,l)} <$ min $FPE_{(k)}$, then the variable X is tentatively said to Granger-cause Y and is retained for further consideration in the Y equation. If min $FPE_{(k,l)} >$ min $FPE_{(k)}$, then the variable X is said not to Granger-cause Y and is tentatively omitted from the Y equation.[6]

Let us assume, as was the case for the models estimated in this paper, that the bivariate equations suggest that both the financial and fiscal variables Granger-cause Y. A determination of the order in which these variables are added to the Y equation must be made. The criterion of Caines, Keng, and Sethi [1], that the variable with the lowest FPE from the bivariate equations is added first, is followed here. Thus the variable with the lowest FPE is added to the Y equation with the lag order from the bivariate equation. The trivariate equation $Y_t = a_0 + a_{11}(L)Y_t + a_{12}(L)X_{1,t} + a_{13}(L)X_{2,t} + e_t$ is estimated where $X_{1,t}$ is the variable with the lowest FPE, $X_{2,t}$ is the remaining variable, and $a_{13}(L)$ is defined analogously to $a_{11}(L)$ and $a_{12}(L)$. $a_{11}(L)$ and $a_{12}(L)$ are fixed at their previously determined orders and the lags in $a_{13}(L)$ are varied over $p = 1, \ldots, m$. The FPEs for the resulting regressions are computed and are defined for lag p, $p = 1, \ldots, m$, as

$$FPE_{(k,l,p)} = [(T+k+l+p+1)/(T-k-l-p-1)][SSR_{(k,l,p)}/T].$$

As before, the lag length that yields the minimum FPE is selected as the lag order for that variable. The FPE is compared to the FPE from the bivariate equation containing the variable with the lowest FPE. If min $FPE_{(k,l,p)} <$ min $FPE_{(k,l)}$, then the variable is tentatively said to Granger-cause Y and is retained for further consideration. If, however, min $FPE_{(k,l,p)} >$ min $FPE_{(k,l)}$, then the variable is tentatively omitted from the Y equation.

6. The Granger-causality statements are tentative at this stage since we are interested in specifying a system and all causality statements must be checked within the context of the system. The results of Nelson-Schwert [20] are again relevant.

Similar procedures are used in specifying the other equations in each model. When the three equations for each model are tentatively specified, they are combined to form a system. One potential problem—contemporaneous relationships among the variables of a system—has been ignored to this point. Following Hsiao [12] it is assumed that any contemporaneous relationships are reflected in the correlation of error terms across the system's equations. Based upon this assumption, full-information maximum likelihood (FIML) is used to estimate the system. The specification of each model is checked by over-and under-fitting the system, estimating the modified systems by FIML, and then carrying out likelihood ratio tests of the adequacy of the specified system against each proposed modification. The likelihood ratio statistics are computed as $-2\ \log(L^c/L^u)$ where L^c is the maximized likelihood of the constrained system (the modified system for underfits but the specified system for overfits) and L^u is the maximized likelihood of the unconstrained system (the specified system for underfits and the modified system for overfits). This statistic asymptotically follows a chi-square distribution with n degrees of freedom, where n is the number of imposed constraints.

III. The Models

The trivariate models estimated using quarterly data for nominal *GNP*, a financial aggregate, and high employment government expenditures are presented in this section. The financial aggregates employed include the aggregates for which the Federal Reserve specified growth rate ranges in the 1970s (*M1, M2, M3*, and bank credit) as well as the liquid asset measure, *L*, and two other credit measures (the debt proxy and total nonfinancial liabilities). The monetary aggregates are generally familiar, and require no elaboration. The credit aggregates range from narrow to broad. The narrowest is bank credit (*BC*), conceptually similar to bank loans and investments and measured on the asset side of bank balance sheets, which includes commercial bank holdings of Federal and state and local obligations and corporate securities as well as total bank loans. The latter include bank holdings of mortgages, consumer credit, open market paper, commercial loans, loans to other financial institutions and loans to foreign banks. The debt proxy (*DP*), an aggregate proposed by Henry Kaufman [14;15] and subsequently analyzed by, among others, Davis [2], includes holdings measured on the asset side of the balance sheets of private domestic nonfinancial investors. These assets include currency, checkable deposits, large and small time deposits, money fund deposits, security RPs, Federal government securities, state and local obligations, and corporate and foreign bonds. *L* is examined since it is closely related to the debt proxy. The broadest aggregate is total nonfinancial liabilities (*TNL*), apparently originally proposed by Friedman [7]. *TNL* represents total credit market funds raised by nonfinancial U.S. borrowers and includes not only private domestic debt (such as corporate bonds, mortgages, consumer credit and open market paper) but also state, local and Federal obligations.

We note two features with respect to these credit aggregates. First, both bank credit and the debt proxy measure items on the asset side of the balance sheets of various agents, while total nonfinancial liabilities are, in fact, measured by the liabilities of particular sectors. Nonetheless, for our purposes, we will refer to these aggregates as "credit" aggregates since, for instance, bank credit represents a credit extended by banks to various borrowers. Second, it is interesting that in 1983 the Federal Reserve has apparently dropped bank credit as an "associated" policy target and replaced it with the total non-financial liabilities aggregate employed here [26]. We include bank credit in our analysis since it was a policy target over the period of investigation and thus provides an interesting contrast with the broader credit measures.

As noted in section II, specification and estimation of the VARs requires stationary data, and it was found that the second difference of logs of each variable yielded stationarity. Intuitively, second differences are reasonable when one considers the generally accelerating inflation during the period under investigation, 1959:1-1979:4. Formally, the second-difference formulation was arrived at, following Hsiao, by regressing the variables transformed in this manner on a constant and time, correcting for serial correlation when necessary. In no case was the coefficient on time significant. However, when first differences of logs of the variables were regressed on a constant and time, the coefficients on time were uniformly statistically significant. Based on these results, second differences of logs were employed.

Using the estimation procedure outlined in section II, the VAR model for the *M1* system, estimated over the 1959:1-1979:4 period, is

$$\begin{bmatrix} Y_t \\ M1_t \\ EHE_t \end{bmatrix} = \begin{bmatrix} a_{11}^{10}(L) & a_{12}^{6}(L) & a_{13}^{1}(L) \\ a_{21}^{9}(L) & a_{22}^{4}(L) & 0 \\ 0 & 0 & a_{33}^{3}(L) \end{bmatrix} \begin{bmatrix} Y_t \\ M1_t \\ EHE_t \end{bmatrix} + \begin{bmatrix} a_1 \\ a_2 \\ a_3 \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \end{bmatrix} \quad (1)$$

where Y_t is the second difference of the log of nominal income, *M1* is the second difference of the log of the narrow money stock, *EHE* is the second difference of high employment spending, a_{ij}^k represents the k lag coefficients on variable j in equation i, the a_i are constants, and the e_i are error terms. As Sims [22] points out, because of the reduced-form nature of the model it is difficult to interpret the individual autoregressive coefficients; these coefficients are not presented here but are available on request.

Before discussing the economic implications of system (1), it is useful to justify the adequacy of the model. Model adequacy is first tested by under- and over-fitting the system and then conducting likelihood ratio tests of the modified systems against system (1). These tests are presented in Table I and are interpreted in the following manner: hypothesis (a) indicates that the lags on Y in the Y and *M1* equations have been lengthened by two quarters relative to those in system (1) while hypothesis (e) indicates that the lags on *M1* in the Y and *M1* equations have been shortened by two quarters relative to system (1).[7] As is evident from tests (a)-(c) in Table I, extending the lags in the system on the various variables never generates a significant improvement in the system. Alternately, hypotheses (d)-(f) in the table suggest that shorter lags on the respective variables are inappropriate. Also reported in Table I are the results of testing various zero restrictions in equation system (1). First, in tests (g)-(i) we have imposed zero restrictions on the non-zero off-

7. The over- and under-fits reported in Tables I and II are obviously only a subset of the potential over- and under-fits that could be tried. However, a limited computing budget restricted the modified systems considered. Additionally, we feel that the over- and under-fits reported in Tables I and II are sufficient to establish the adequacy of models (1)-(7).

Table I. Tests of Model Adequacy, Equation System 1

Hypothesis	Chi-Square Statistic[a]
a. a_{11}^{12}, a_{21}^{11}	9.35
b. a_{13}^{3}, a_{33}^{5}	1.08
c. a_{12}^{8}, a_{22}^{6}	3.47
d. a_{11}^{8}, a_{21}^{7}	19.93*
e. a_{12}^{4}, a_{22}^{2}	32.68*
f. a_{33}^{1}	19.86*
g. $a_{12} = 0$	26.93*
h. $a_{13} = 0$	16.99*
i. $a_{21} = 0$	20.14*
j. $a_{31}^{4} \neq 0$	1.05
k. $a_{32}^{4} \neq 0$	6.04
l. $a_{23}^{4} \neq 0$	1.30

a. * = significant at .005 level.

diagonal elements and, at the usual significance levels, rejected restrictions of this nature. Second, in tests (j)-(l), we eased existing zero restrictions and found that the zero restrictions were always appropriate. Thus, system (1) appears to be an adequate VAR representation of the three variables in question.

A number of substantive points about system (1) may be made. First, both money and high employment spending cause nominal income, results evident from the facts that $a_{12}(L)$ and $a_{13}(L)$ are significantly non-zero. Second, "reverse" causation from income to *M1* exists. This feedback may be due to shifts out of *M1* in response to changes in interest rates stemming from prior movements in income [12]. Whatever the reason may be, the result suggests that a traditional "reduced-form" regression of income on narrow money and high employment spending may be subject to simultaneous equations bias. Third, neither income nor money causes *EHE*, so that fiscal policy, as measured here, appears to be exogenous.[8]

The specifications of the trivariate models for the remaining monetary aggregates for the period 1959:1-1979:4 are presented in Table II. In order to conserve space, the results

8. These results are at odds with those presented by Mehra-Spencer [16] who concluded that it was fiscal policy, not monetary policy, which was endogenous. This difference with the current study may be explained by the use of different sample periods, our use of growth rate transformations versus Mehra-Spencer's use of several different transformations of the first differences of levels of the data, our use of lag lengths determined by the data rather than Mehra-Spencer's use of four and eight quarter lags in their Granger-Sargent tests and four quarter leads and four and eight quarter lags in their Sims' tests, and our focus upon a system of reduced-form equations rather than a single equation.

Table II. Trivariate Models for Monetary and Credit Aggregates

System	Non-Zero Elements[a]
(2) *Y*, *M2*, *EHE*	a_{11}^{10}, a_{12}^{14}, a_{13}^{1}, a_{22}^{12}, a_{33}^{3}
(3) *Y*, *M3*, *EHE*	a_{11}^{10}, a_{12}^{10}, a_{13}^{1}, a_{22}^{3}, a_{33}^{3}
(4) *Y*, *L*, *EHE*	a_{11}^{10}, a_{12}^{5}, a_{13}^{1}, a_{22}^{1}, a_{33}^{3}
(5) *Y*, *BC*, *EHE*	a_{11}^{10}, a_{12}^{3}, a_{13}^{1}, a_{22}^{1}, a_{33}^{3}
(6) *Y*, *DP*, *EHE*	a_{11}^{10}, a_{12}^{3}, a_{13}^{1}, a_{22}^{1}, a_{33}^{3}
(7) *Y*, *TNL*, *EHE*	a_{11}^{10}, a_{12}^{7}, a_{13}^{1}, a_{21}^{7}, a_{22}^{8}, a_{33}^{3}

a. a_{ij}^{k} represents the k lag coefficients on variable j in equation i. Thus in system (2) a_{12}^{14} indicates that the lag on *M2* in the nominal income equation is 14 quarters.

of the tests of model adequacy similar to those in Table I are not presented here but are available on request from the authors. Based on these tests, each model is judged adequate.

Of particular interest in systems (2)-(4) are the facts that the broader monetary aggregates and high employment spending are both free of reverse causation from income (and each other as well). The absence of feedback from income to these broader aggregates is in contrast to the feedback from income to *M1* and may be explained by the fact that much of the shift out of *M1* is into the components of the broader aggregates so that on balance the broader aggregates are unaffected by prior movements in income.

Two further factors are of note with respect to systems (2)-(4). First, even though single-equation income regressions can produce historical estimates of unbiased policy multipliers for the broader monetary aggregates, the respective three-equation systems are useful for forecasting purposes since each system implies univariate forecasts for the exogeneous policy variables. Second, a well-known, desirable property of a good target of policy is that it should be free from feedback from non-policy variables. If not, it is uncertain whether movements in the variable are the results of policy actions or of economic forces beyond the authorities' control. The absence of feedback from income to the broader monetary aggregates, in contrast to *M1*, suggests the potential usefulness of these aggregates.

We next estimated over 1959:1-1979:4 analogous systems employing credit aggregates rather than monetary aggregates as the relevant financial variable. The resulting systems (5)-(7) are presented in Table II. The adequacy of these models is, as with the monetary models, justified by tests similar to those reported in Table I. Each model is judged adequate, and the test results are available upon request.

We notice in systems (5) and (6) that use of either bank credit or Kaufman's debt proxy in a St. Louis-type reduced-form does not generate reverse causation problems. Rather, in each of these cases, the financial aggregate and high employment spending follow univariate processes. In system (7), however, we do find that income Granger-causes total net liabilities, as well as the converse. The presence of reverse causation in this broadest credit aggregate may well reflect the responsiveness of total credit demands with the business cycle, particularly in view of the fact that *TNL* is the only one of the credit aggregates which is measured on the liability side of the balance sheets of various agents.

The results of this section uniformly indicate that, with appropriate stationarity transformations, high employment spending is not the source of simultaneity bias in reduced-form St. Louis-type regressions. Rather, when such reverse causation exists, it involves the financial aggregate employed. For the seven financial series used here, only the narrow money and broadest credit measures involve feedback from income.

IV. Further Implications

The estimated systems can be used to analyze the effects of changes in the growth rates of the various financial aggregates and high-employment expenditures on the growth of nominal income. In particular, we first present reduced-form monetary and fiscal policy multipliers for those models in which the financial aggregates are not subject to reverse causation from income; multipliers associated with aggregates (in our case, *M1* and *TNL*) subject to such reverse causation are not presented since their endogeneity raises the question of whether these aggregates should be used as policy targets. Second, we investigate the forecasting performance for each system. Both exercises are performed after transforming all variables to first differences of logs, a transformation employed since the focus in the literature is on the effects of the rates of growth of the financial aggregates and high employment expenditures.

The dynamic multipliers for changes in the financial aggregates and high-employment expenditures are derived from dynamic simulations. For each system, a base dynamic simulation using historical data for all variables was first performed. Then a dynamic simulation was run with the growth rate of the financial aggregate increased by 1 percent above its historical values; historical data were used for the other variables in the system. Another dynamic simulation was run for each system with the growth rate of high-employment expenditures raised by 1 percent above its historical values; again, historical data were used for the other variables in the system. The difference between the base simulations and the policy simulations provide an estimate of dynamic multipliers for the financial aggregates and high-employment expenditures.[9]

The dynamic multipliers are presented in Table III. From part A of this table, we see that a 1 percent increase in the growth rate of *M2* raises nominal GNP growth by 0.9 percent at the end of a year; the effect declines somewhat over years two through four. However, the long-run effect (effect in quarter 20) is an approximate 1 percent increase in nominal GNP growth. The long-run effects of increases in the other financial aggregates is less than for *M2*, although the effect at the end of the fourth quarter for *L* and *DP* are similar to the four quarter effect of *M2*. Increases in the growth rates of *M3* and *BC* have the smallest effects on nominal GNP growth. The effect at the end of the fourth quarter is

9. An objection based on the Lucas critique might be raised about the calculation of these multipliers. Specifically, it might be argued that the sustained 1 percent increase in the growth of the financial aggregates above their historical values represents a regime change which in turn leads to a change in the model parameters. This change in turn renders the multiplier estimates suspect. Sims [24,122] argues that this critique of policymaking is an example of the general statistical principle that "... statistical models are likely to become unreliable when extrapolated to make predictions for conditions far outside the range experienced in the sample." Certainly the experiments we consider are not far outside the historical ranges for growth of the financial aggregates. Finally, past evaluations of potential targets for monetary policy have examined the closeness of the relationship of alternative aggregates to *GNP* by computing and analyzing multipliers of the sort we compute

Table III. Dynamic Multipliers for Nominal GNP Growth

Financial Aggregates[a]						High-Employment Expenditures[b]					
	System						System				
Period	*M2*	*M3*	*L*	*BC*	*DP*	Period	*M2*	*M3*	*L*	*BC*	*DP*
4	.93	.48	.96	.39	.97	4	.01	.01	.01	.02	.02
8	.74	.45	.36	.09	.54	8	.02	.02	.03	.02	.03
16	.74	.38	.52	.20	.59	16	.02	.02	.02	.02	.03
20	.98	.52	.64	.18	.69	20	.01	.02	.03	.02	.04

a. Part A shows the response of the growth rate of nominal GNP to a sustained 1 percent increase in the growth rate of the financial aggregate.

b. Part B shows the response of the growth rate of nominal GNP to a sustained 1 percent increase in the growth rate of high-employment expenditures.

0.5 percent for *M3* and 0.4 for *BC* while the long-run effects are 0.5 percent for *M3* and only 0.2 percent for *BC*. The long-run effects for *L* and *DP* are greater than for *M3* and *BC*, 0.6 percent and 0.7 percent, respectively, but are still below the long-run effects for *M2*. Thus, in terms of magnitudes of multipliers, *M2* has the largest effect over time, followed by *DP* and *L*. Furthermore, the size of the multiplier for *M2* appears less variable over time than the multiplier for the other financial aggregates.

The effect of a 1 percent increase in the growth rate of high employment expenditures is very similar across the different systems. Even though *EHE* is exogenous in each system the effects are not exactly the same since the lags on *Y* and the financial aggregates differ across systems. From part B of Table III, we see that an increase in high-employment expenditure growth has virtually no effect in either the short-run or the long run.

The various systems were also compared on the basis of out-of-sample forecasting performance. The systems were dynamically simulated over the six quarter period beginning 1980:1. The results are presented in Table IV, Part A. We note that the root-mean-square errors for nominal *GNP* growth are very similar for the alternative systems, with the *M2* and *DP* systems producing slightly lower root-mean-square errors. The forecasts can be further analyzed by examining the decomposition of the mean-square errors into bias, regression, and residual proportions. The bias and regression proportions tend to zero for an optimal predictor. The lowest bias proportion is found for the *M2* system, followed by the *DP* system. However, the lowest regression proportion is generated by the *DP* system. On the whole, there are no substantial differences across systems in terms of root-mean-square error performance, although the *M2* and *DP* systems generate somewhat lower root-mean error performance and the bias proportion for these systems is lower than for the other systems.

For purposes of comparison, standard "St. Louis" equations were estimated for the financial aggregates found to be exogenous. The polynomial lag, growth rate specification of Seaks-Allen [25]—a fourth degree polynomial with both end point constraints and current and four lagged values for both the financial aggregate and high-employment

Table IV. Out-of-Sample Forecasting Performance[a]

	VAR Systems				St. Louis Equations			
	Root-Mean-Square Error for Nominal GNP Growth	Decomposition of Mean-Square Error			Root-Mean-Square Error for Nominal GNP Growth	Decomposition of Mean-Square Error		
System		Bias Proportion	Regression Proportion	Residual Proportion		Bias Proportion	Regression Proportion	Residual Proportion
M2	.013	.003	.028	.969	.014	.0004	.886	.114
M3	.015	.014	.005	.981	.014	.0004	.716	.284
L	.016	.041	.021	.939	.013	.013	.241	.746
BC	.014	.011	.097	.892	.013	.025	.358	.617
DP	.013	.005	.001	.994	.012	.002	.342	.656

a. The simulations were performed over the six quarter period beginning in 1980:1 and ending in 1981:2.

expenditures—was used. Out-of-sample forecasts for the six quarter period beginning in 1980:1 were computed and are presented in part B of Table IV. The root-mean-square errors are comparable to those from the VAR systems, as are the bias components of the forecasts. Indeed, these measures of forecast performance are typically somewhat lower than for the VAR systems. However, the quality of the St. Louis forecasts in terms of the regression and residual proportions of the mean-square error is much poorer than for the VAR systems. Thus in terms of the quality of the forecasts the VAR forecasts, on balance, appear superior to the single equation forecasts.

V. Summary

The objective of this paper has been to investigate the empirical characteristics of the relationship among nominal GNP, high-employment expenditures, and various financial aggregates in terms of the characteristics of a good intermediate target. The choice of variables is reminiscent of the standard "St. Louis" equation; however, the analysis is conducted within the framework of recent advances in reduced-form methodology, *viz.*, vector autoregressions.

The financial aggregates included in the discussion are the standard monetary aggregates (*M1*, *M2*, *M3*, and *L*) and a set of credit aggregates (bank credit, the debt proxy, and total nonfinancial liabilities). Bank credit, *M1*, *M2*, and *M3* were analyzed in light of the explicit roles these aggregates have played, over the past decade, as intermediate policy targets. As discussed by Fackler-Silver [5], the remaining aggregates were examined in view of the fact that recent financial innovations have caused the Federal Reserve to focus on broader aggregates, on both sides of the balance sheet, in the formulation and implementation of policy. In particular, the Federal Reserve has recently announced growth rate targets for total nonfinancial liabilities.

We find that two of the Federal Reserve's current intermediate targets—*M1* and total nonfinancial liabilities—are subject to feedback from income and thus may not serve well as intermediate targets. The other aggregates are free of this feedback and therefore possess one of the characteristics of a good target. Estimated policy multipliers for the remaining aggregates suggest that fiscal policy has virtually no impact on income but that the influence of the financial aggregates, with the exception of bank credit, is substantial. *M2*, one of the current intermediate targets, has the largest impact on income. Finally, it appears that the quality of the forecasts from the vector autoregressive systems is better than forecasts from single reduced-form equations.

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