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IDENTIFICATION AND THE LIQUIDITY EFFECT OF A MONENTARY POLICY SHOCK

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ABSTRACT

Conventional wisdom holds that unanticipated expansionary monetary policy shocks cause transient but persistent decreases in real and nominal interest rates. However a number of econometric studies argue that the evidence favors the opposite view, namely that these shocks actually raise, rather than lower, short term interest rates. We show that this conclusion is not robust to the measure of monetary aggregate used or to the assumptions made to identify monetary policy disturbances. For example, when our analysis is done using non borrowed reserves, we find strong evidence in favor of the conventional view. Existing challenges to the conventional view lack credibility not just because of their fragility. They are based upon measures of policy disturbances which generate seemingly implausible implications about things other than interest rates.

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1. Introduction

Conventional wisdom holds that unanticipated expansionary monetary policy shocks cause transient but persistent decreases in real and nominal interest rates. However, there is virtually no formal econometric evidence in the literature to support this contention. (See Reichenstein 1987 for a review of the literature.) Indeed, many empirical studies have concluded that the data support the opposite view, namely, that unexpected shocks to monetary policy actually raise, rather than lower, short–term interest rates. (See for example Mishkin 1981, 1982 and R. King 1991.) This finding typically is rationalized as reflecting the weakness of liquidity effects associated with expansionary monetary policy and the relative strength of expected inflation effects on nominal interest rates. (See, for example, Mishkin 1981, 1982 or S. King 1983.) We argue here that an analysis of the data which pays particular attention to the problem of measuring monetary policy shocks reveals substantial support for the conventional wisdom. Existing results in the literature which purport to cast doubt on that wisdom are not robust to plausible changes in identifying assumptions or to alternative ways of measuring money.

Analysts of monetary policy must confront two problems. One is just which measure of money to use. After all, we have at our disposal a plethora of measures ranging from narrow, direct measures of open market operations like nonborrowed reserves (NBR) to relatively broad aggregates like M2. The choice of money measure has important implications for inference because different monetary aggregates interact in very different ways with short–term interest rates. The other problem is which set of identifying assumptions to adopt to measure the exogenous component (if any) of changes in monetary policy. Without such assumptions, causal inference is simply not possible. In practice, these two choices – of monetary aggregate and identifying assumptions – cannot be viewed as distinct problems because there is no reason to believe that any given set of identifying assumptions will be equally appropriate across different measures of money.

These difficulties notwithstanding, empirical work on the liquidity effects of monetary policy almost always uses high–order monetary aggregates like M1.[1] This choice of monetary aggregate is usually coupled with the identifying assumption that the monetary policy disturbance corresponds to the statistical innovation in money. Put differently, innovations to objects like M1 are entirely attributed to the actions of the monetary authority.[2] In our view, an alternative assumption is at least as plausible, namely, that innovations to NBR capture Fed policy shocks, while innovations to broader aggregates confound many other shocks, in addition to policy shocks.[3]

In contrast to the existing literature, in this paper we use various measures of money and different identifying assumptions to measure unanticipated shocks to monetary policy.[4] For each measure of money considered (NBR, the monetary base, M0, and M1), we engage in a specification search across the elements of two classes of identifying assumptions. The first class, which we call *M–rules*, is defined by the assumption that unanticipated changes in monetary policy can be measured by some orthogonalized component of the innovation to the monetary aggregate. Each such component corresponds to a different assumption regarding the variables included in the contemporaneous portion of the Federal Reserve Open Market Committee's (FOMC's) reaction function for setting

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[1] For example, Mishkin (1981, 1982), Cochrane (1989), and R. King (1991) use M1 while Melvin (1983) uses M2, and Reichenstein (1987) uses both M1 and M2. Indeed, Reichenstein's (1987) review of this literature does not contain even one reference to a study which uses NBR or even M0 as the measure of money. Recently, Gordon and Leeper (1991) use M0 in their analysis while Terhan (1991) uses weekly data on NBR.

[2] This assumption is implicit in studies that regress the interest rate on the unanticipated change in money, where the latter is the residual from an equation in which the time t value of a monetary aggregate is regressed against time t–1 variables. Such a procedure is asymptotically equivalent to running a vector autoregression which contains, among other things, money and interest rates, and then calculating the moving average representation implied by the assumption that the time t disturbance to money is orthogonal to innovations in the other variables in the system. In Sims' (1986) terminology, this identification scheme amounts to adopting a Wold causal chain in which money is placed first in the ordering.

[3] NBR is total reserves, less total borrowings of depository institutions from the Federal Reserve. (See Table 1.20, on page A12 of U.S., Federal Reserve Board 1991a.) NBR is the monetary aggregate most closely controllable by the FOMC. Broader aggregates like M1 and M2 are less closely controllable, because the non–NBR component in these aggregates is observed with a lag.

[4] Sims (1986) looks at different identifying assumptions conditional on using a particular monetary aggregate, M1, in his analysis.

the monetary aggregate. While quite natural, this class of identifying assumptions is by no means uncontroversial. McCallum (1983), Sims (1986, 1991), and Bernanke and Blinder (1990) have argued that – at least for high–order aggregates like M1 and M2 – unanticipated shocks to monetary policy are best measured as the innovation in the federal funds rate. For this reason, we consider a second class of identifying restrictions which is defined by the assumption that unanticipated changes in monetary policy can be measured by some orthogonalized component of the innovation to the federal funds rate. Each such component corresponds to a different assumption about the variables included in the contemporaneous portion of the FOMC's rule for setting the federal funds rate. We refer to this class of policy rules as *R–rules.*[5]

Our empirical analysis reveals that inference about the effects of monetary policy on interest rates hinges critically on both the identifying assumptions exploited and the measure of money used. Certainly we have found combinations of monetary aggregate and identification schemes which together generate challenges to the conventional view. But in every such instance the associated measure of unanticipated shocks to monetary policy generates seemingly implausible implications about things other than interest rates. When these combinations of monetary aggregates and identification schemes are eliminated, the remaining combinations all yield results which strongly support the conventional view. Specifically, using exactly identified vector autoregressions, we find that when identifying assumptions corresponding to M–rules are coupled with either M0 or M1, unanticipated changes in monetary policy generate increases in the federal funds rate. However, so identified, unanticipated expansionary monetary policy shocks generate sharp, persistent declines in aggregate output. Indeed, our point estimates indicate that, for M1, real gross

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[5] A simple way to understand the difference between M— and R— rules is to consider the extreme case in which each fails to feed back on the contemporaneous value of any other variable. This version of the M–rule corresponds to a perfectly interest–inelastic short–run money supply rule, while this version of the R–rule corresponds to a perfectly elastic short–run money supply rule in which all shocks to money demand are completely accommodated.

national product (GNP) falls for over nine years. One cannot accept the interest rate implications of these measures of monetary policy shocks without also accepting these seemingly implausible aggregate output implications.[6]

Inference with M0 and M1 is greatly affected by moving to the class of identifying restrictions corresponding to R–rules. Under these circumstances, we find that unanticipated changes in monetary policy generate sharp, persistent declines in the federal funds rate. Moreover, when measured in this way, unanticipated expansionary monetary policy generates persistent increases in aggregate real output. We infer that if one insists on using high–order monetary aggregates to study the effects of monetary policy on interest rates, this class of identifying restrictions is preferable to the class of M–rules.

Interestingly, in sharp contrast to results based on M0 or M1, inference about the effects of monetary policy on interest rates is very robust when the aggregate NBR is used in the analysis. Regardless of whether we work with M– or R–rules, regardless of whether we work with monthly or quarterly data, and regardless of which postwar sample period we work with, the same result emerges. Unanticipated expansionary policy shocks drive down short–term interest rates for substantial periods of time. Measured in this way, expansionary monetary policy shocks also generate increases in real GNP. It is hard to imagine reconciling these findings with models that are empirically plausible and yet do not incorporate quantitatively important liquidity effects.[7] Building on earlier contributions by Grossman and Weiss (1983) and Rotemberg (1984), recent work by Lucas (1990), Fuerst (1990), Baxter, Fischer, R. King and Rouwenhorst (1990), Christiano (1991), and

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[6] There do exist economic models that can rationalize a fall in output following a money shock. (See Cooley and Hansen 1989, Christiano 1991 and Christiano and Eichenbaum 1991.) However, we suspect that plausibly parameterized versions of these models will have difficulty accounting for the magnitude and persistence of the fall in output.

[7] For example, one could construct a model in which the fundamental shocks driving the business cycle are non–monetary and have the effect that, in equilibrium, there is a positive association between output growth and the interest rate. If such a model incorporated a monetary policy of "leaning against the wind" — tightening money in a boom and easing in a recession — then it would imply a negative association between money and interest rates, even if the model had no liquidity effect. However, such a model would have the counterfactual implication that output growth and the interest rate are positively correlated.

Christiano and Eichenbaum (1991) has stressed the importance of liquidity effects for explaining the comovements between interest rates and monetary aggregates. We view our results as being very supportive of that work.

The paper is organized as follows. In section 2 we present some basic facts about the dynamic (unconditional) correlations between different monetary aggregates and the federal funds rate. Three key facts emerge there. First, the federal funds rate displays a sharp, robust, negative correlation with NBR. This negative correlation is masked by moving to higher–order monetary aggregates. Second, once we control for the behavior of borrowed reserves (BR), M0 behaves much like NBR. Third, the federal funds rate displays sharp, persistent, negative comovements with real GNP. The dynamic correlations between these two time series are estimated very precisely and are very robust to sample period selection. While highly suggestive, the results cannot be taken as evidence of any specific causal mechanism. In particular, they cannot be used to formally infer that unanticipated expansionary monetary policy disturbances cause interest rates to fall and aggregate real output to rise. Such conclusions necessarily rely on theoretical restrictions which enable the analyst to identify the exogenous component of monetary policy disturbances. This issue is addressed in section 3, where we abandon the sharp distinction between theory and measurement and use vector autoregressions to interpret the relationship between money and interest rates. Finally, in section 4 we summarize our findings.

2. Some Basic Facts: The Dynamic Correlations Between Money, the Federal Funds Rate, and Real Output

In this section we report some basic facts about the dynamic correlations between different measures of money, the nominal federal funds rate and aggregate output. Our primary findings are that the nominal federal funds rate displays strong negative co–movements with different measures of the growth rate of money and of aggregate real output. The relationship between the federal funds rate and aggregate real output, as summarized by their dynamic correlations, is estimated very precisely and is very robust across different sampling periods as well as different stationary–inducing transformations of the data. The negative relationship between the federal funds rate and the growth rate of money is most pronounced and most stable when NBR is used as the measure of money.

In presenting these findings, we are mindful of the obvious caveat that correlations do not imply causality. Still, the results in this section serve at least three useful functions. First, these correlations represent important moments of the data that any business cycle model ought to be consistent with. Second, the correlations suggest that monetary business cycle models that display significant, persistent liquidity effects will be useful for interpreting the data. Finally, they provide a useful background for the vector autoregression analysis of section 3.

2.1 Choosing Measures of Short Term Interest Rates and Money

Figure 1 displays the nominal federal funds rate (FF), the six–month nominal commercial paper rate, and the three–month nominal Treasury bill rate over the sample period from the first quarter of 1959 to the first quarter of 1990 (1959:1 – 1990:1). Notice that these different short term interest rates exhibit similar trends and move together quite closely. In what follows, we display results based on FF for two reasons. First, consistent with Figure 1, our results are not very sensitive to which interest rate we work with. Second, numerous authors have stressed the important role that the federal funds rate plays in monetary policy. (See for example Bernanke 1990, Bernanke and Blinder 1990 and Kuttner and Friedman 1990.) Working with the federal funds rate facilitates comparisons with this literature.

Choosing which measure of money to work with is a much more difficult task. Existing monetary theories of the business cycle are simply too abstract to warrant focusing on any one measure of money. Adding to the difficulty is that different components of any given measure of money often behave in very different ways. Consider, for example, high–powered money, M0, which is the sum of currency in the hands of the public, plus NBR, plus BR. As we shall show, the federal funds rate is negatively correlated with NBR but positively correlated with BR.

In part, the different behavior of NBR and BR simply reflects the institutional reality of how BR is determined. Of particular note is that discount window borrowing is administered under a set of guidelines that is independent of the deliberations of the FOMC. (See Goodfriend and Whelpley 1986 or Stigum 1990.)[8] In contrast, the FOMC directly controls, by open market operations, the level of NBR. From this perspective, NBR seems like a natural measure of money to use in identifying and estimating the effects of shocks to monetary policy. At the same time, we recognize that NBR is not necessarily the best measure of money for assessing the overall empirical plausibility of monetary business cycle models. Consequently, we also present results using M0 and M1. Finally, because of the importance of BR for some of the moments which we discuss, we also display results using BR, M0 less BR, and M1 less BR.

Figures 2a and 2b display seasonally adjusted average quarterly NBR (adjusted to include extended credit) and FF over the sample period 1959:1 to 1990:1. As can be seen, both exhibit a strong positive trend. Other measures of the money supply, such as the level of M0 and M1, also display pronounced trends over this sample period: Consequently, some stationary–inducing transformation of the data must be adopted in order to calculate meaningful statistics. In this section, we work primarily with the filter developed by Hodrick and Prescott (1980). However, we also present results with linearly detrended data

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[8] The basis for lending at the discount window is laid out in the Fed's Regulation A, according to which "Federal reserve credit is available on a short–term basis to a depository institution under such rules as may be prescribed to assist the institution, to the extent appropriate, in meeting temporary requirements for funds, or to cushion more persistent outflows of funds pending an orderly adjustment of the institution's assets and liabilities." (U.S., Federal Reserve Board 1991, sec. 201.3, par. (a)).

and growth rates. The dark lines in Figures 2a and 2b display the Hodrick–Prescott (HP) trend component of NBR and FF, respectively. In Figure 3a we display the HP–filtered versions of NBR and FF. Note the pronounced negative association between these variables. This basic fact is reflected in all of the formal statistics presented in this section. Figure 3b presents the HP filtered versions of FF and real GNP. FF is positively correlated with the contemporaneous level of GNP but negatively correlated with future levels of GNP.

2.2 A Benchmark Scenario

Before discussing our empirical results, we digress to consider the question: What pattern of dynamic correlations would we expect to find in the presence of liquidity effects? A precise answer to this question obviously requires a formal model.[9] Still it seems worthwhile appealing to existing models in the literature to provide some perspective on our reduced–form results. In so doing we assume, as is the case for our measures of money, that money, M_t, is positively correlated over time. Also, for simplicity, we consider a benchmark scenario in which the only shocks are to the money supply.

Consider first the correlation between FF_t and future values of M_t. Suppose that, at time t, there was an unanticipated increase in the money supply. Given a liquidity effect, this would be associated with a decline in FF_t. With M_t positively correlated over time, high values of M_t would be associated with high values of $M_{t+\tau}$, for $\tau > 0$. Other things equal, then, we would expect FF_t to be negatively correlated with future values of M_t, with the exact magnitude of the correlation depending on the size of the liquidity effect and the degree of serial correlation in M_t.

Next consider the correlation between FF_t and past values of M_t. Suppose that at

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[9] Recent examples of such models are provided by Baxter, Fischer, King and Rouwenhorst (1990); Lucas (1990); Fuerst (1990); Christiano (1991); and Christiano and Eichenbaum (1991).

time $t-\tau$, for $\tau > 0$, there was an unanticipated increase in the money supply. This would exert negative pressure on $FF_{t-\tau}$. Suppose that M_t is sufficiently autocorrelated that the initial increase in $M_{t-\tau}$ is associated with higher growth rates in $M_{t-\tau+j}$ for $j \geq 1$. This, we expect, would generate an increases in the anticipated rate of inflation from time $t-\tau+j$, for $j \geq 1$. If the liquidity effect lasted only one period, then the inflation effect would dominate after one period, so that $FF_{t-\tau+j}$, for $j \geq 1$, would rise. Consequently, $FF_{t-\tau+j}$, for $j \geq 1$, would be positively correlated with $M_{t-\tau}$, that is, $\rho(FF_t, M_{t-\tau}) > 0$ for $\tau \geq 1$, where $\rho(\cdot,\cdot)$ denotes the correlation operator. In fact, there is no reason to believe that the liquidity effect lasts for only one period. Suppose instead that it dominated the expected inflation effect for k periods. Then $\rho(FF_t, M_{t-\tau})$ would be negative for $\tau \leq k$, but positive for $\tau > k$. In this sense, k can be thought as measuring the persistence of the liquidity effect.

While useful for pedagogical purposes, the logic of the previous scenario holds only if the sole source of aggregate uncertainty is shocks to the money supply. With other shocks to the system, the dynamic correlation between FF and the stock of money depends, at least in part, on the way the FOMC reacts to the other shocks. For example, shocks that stimulate money demand tend to create a positive association between money and interest rates in an environment where the Fed seeks to smooth nominal interest rates. Still, were the pattern of correlations arising from our benchmark scenario completely absent from the data, we would have no obvious reason to seek evidence for liquidity effects in the context of more complicated, multiple–shock representations of the data. In fact, the actual pattern of correlations is consistent with our benchmark scenario. This provides additional motivation for the analysis of section 3, where we abandon the one–shock premise of the benchmark scenario and analyze the data using exactly identified vector autoregressive representations of the data.

2.3 The Dynamic Correlations Between The Nominal Federal Funds Rate and Money

Figure 4 presents, graphically, our point estimates of $\rho(\mathrm{FF}_t, \mathrm{M}_{t-\tau})$ for $\tau = -8,\ldots,8$, corresponding to three stationary–inducing transformations of the data: HP–filtered (the first column), linear detrending (the second column), and growth rates (the third column). The three rows contain results pertaining to NBR, M0, and M1 as the measure of money. All correlations are based on variables which have been logged prior to the stationary–inducing transformation. Figure 5 presents the analog results for BR, M0 less BR, and M1 less BR. The solid lines in Figures 4–5 denote point estimates of the correlations in question, along with a two–standard–deviation band, given by the dashed lines. (Standard errors were computed using a generalized method of moments procedure.)

We begin by discussing results obtained with HP–filtered versions of the data. Consider first the estimated values for $\rho(\mathrm{FF}_t, \mathrm{M}_{t-\tau})$ for $\tau = -8,\ldots,8$, when NBR is used as the measure of money. Three findings here are notable. First, there is a strong, statistically significant, negative contemporaneous correlation (equal to –0.39) between FF_t and M_t. This is consistent with the impression conveyed by Figure 3. Second, FF_t is negatively correlated with leads and lags of M_t up to one year. Third, FF_t is positively correlated with $\mathrm{M}_{t-\tau}$ for $\tau > 4$. These three findings are consistent with the benchmark scenario in which a liquidity effect of a monetary policy shock dominates the anticipated inflation effect for a period as long as a year.[10]

The point estimates obtained with the broader measures of money display a somewhat different pattern. As with NBR, our point estimate of $\rho(\mathrm{FF}_t,\mathrm{M}_t)$ is negative when M1 is used as the measure of money (–0.02). However, unlike with NBR, the maximal negative correlation occurs at $\tau = -3$ rather than $\tau = 0$ ($\rho(\mathrm{FF}_t, \mathrm{M}_{t+3}) = -0.33$). With M1, as with NBR, FF_t is negatively correlated with current and future values of money. But unlike the results obtained with NBR, with M1, FF_t is positively correlated with all past values of money. Interpreted from the perspective of the benchmark scenario,

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[10] In order to check the robustness of these results, we redid our calculations using the nominal three–month Treasury bill as our measure of the interest rate. This change has virtually no impact on our conclusions.

these results are consistent with the existence of a strong liquidity effect, but one which is less persistent than the effect observed with NBR.

The only measure of money with which FF_t displays a positive contemporaneous correlation is M0 ($\rho(FF_t,M_t) = 0.25$). Even here, though, FF_t is negatively correlated with $M0_{t+\tau}$, for $\tau \geq 1$. As it turns out, the negative contemporaneous correlation between M0 and FF is attributable entirely to the BR component of M0. As Figure 5 reveals, HP–filtered BR displays a very strong positive contemporaneous correlation with HP filtered FF_t. Presumably this reflects the incentive of banks to increase BR in response to an increase in FF and the practice of accommodating transient increases in bank demand for reserves through discount window lending. (See Thomas 1982, Goodfriend 1983, Goodfriend and Whelpley 1986, and Stigum 1990.) Notice that when BR is subtracted from M0, the resulting dynamic correlations are very similar to those between FF and NBR. As might be expected, subtracting BR from M1 strengthens the negative correlation between FF_t and current and future values of M_t but dampens the positive correlation between FF_t and past values of M_t.

Consider next the results of working with growth rates of the data. Three features of these results are worth noting. For every measure of money, there is a strong negative correlation between FF_t and M_t, regardless of whether or not we control for BR. In addition, the growth rate of FF_t is negatively correlated with future values of the growth rates in NBR, M0, and M1. Finally, FF_t is positively correlated with past values of growth rates in M0 and M1. This tendency is less pronounced with NBR.

Consider next our results with linearly detrended data. A number of comments are in order here. First, with this stationary–inducing transformation, FF is negatively correlated with current levels of the money supply and all its leads and lags ($\tau = -8,...0,...,+8$) regardless of whether NBR, or M0, or M1 is used. Second, with M0 and M1, the shape of the correlation functions using HP–filtered data and linearly detrended data are quite different. In our view, this reflects the dubious validity of the assumption that M0 and M1 are well represented as trend stationary processes with a constant trend over the sample period as a whole. (See Stock and Watson 1989b, who argue that, over this period, the money growth rate has an upward trend.)

In order to assess the sensitivity of our results to sample period selection, we redid our analysis allowing for a break in the data at 1979:3. Figures 6, 7, and 8 present our results for NBR, M0, and M1, respectively, for the sampling intervals 1959:1–1990:1, 1959:1–1979:3, and 1979:4–1990:1. From Figure 6 we see that, despite some differences, the results obtained with NBR are very stable across sample periods. However, Figure 7 reveals considerable sample period sensitivity with M0. As a rule, the post–1979 and full sample correlation functions are similar, at least when we work with HP–filtered data or growth rates. However, the pre– and post–1979 periods results are quite different – so different, in fact, that inference regarding the plausibility of the benchmark scenario is substantially influenced by sample period selection. For the post–1979 period, the dynamic correlations appear to be entirely consistent with that scenario. For example, working with HP filtered data, we find that $\rho(\mathrm{FF}_t, \mathrm{M0}_{t-\tau}) < 0$ for $\tau < 1$. In contrast, the pre–1979 results seem difficult to reconcile with the benchmark scenario. Finally, Figure 8 reveals that sample period sensitivity with M1 is intermediate to the two polar cases of NBR and M0.

We summarize our findings for the correlations between money and interest rates as follows. First, when NBR is used, our results are consistent with what is to be expected from the benchmark scenario. This is true regardless of which stationary–inducing transformation or sample period is used. Second, the results with M0 and M1 seem to depend more sensitively on the sample period used. For the post–1979 period the dynamic correlations of these aggregates seem to accord well with the benchmark scenario. Third, we find that the time series properties of BR and NBR are very different. Simply adding the two when working with monetary aggregates like M0 obscures fundamental differences in the ways that these two types of reserves interact with interest rates.

We conclude this section by briefly discussing the dynamic correlations between FF and per capita real GNP (Y). These are summarized in Figure 9, which displays our point estimates of the correlations for three stationary–inducing transformations of the data and three sample periods. Notice that, while the contemporaneous correlation between FF_t and Y_t is positive, FF_t displays a strong, sharp, negative correlation with future values of Y_t. This is true independent of which stationary–inducing transformation is used or which sample period is investigated. Interestingly, the correlation function of FF_t and Y_t seems to be estimated much more precisely than the correlation functions between money and interest rates. These results are consistent with findings by Kuttner and Friedman (1990), Bernanke (1990), Bernanke and Blinder (1990), and Stock and Watson (1989b). They show that the nominal federal funds rate is an excellent statistical predictor of real GNP, with positive movements in FF preceding declines in real GNP.

3. Vector Autoregressions and the Liquidity Effect

We have documented the existence of a strong negative correlation between different measures of the growth rates of the money supply and GNP with the federal funds rate. However, while highly suggestive, these correlations cannot, in and of themselves, be taken as evidence that unanticipated expansionary monetary policy disturbances drive interest rates down and aggregate output up. At a minimum, providing such evidence requires identifying assumptions that are sufficiently strong to isolate a measure of monetary policy disturbances. As it turns out, inference regarding the effects of monetary policy on interest rates hinges critically on two things: the identifying assumptions used to obtain measures of unanticipated shocks to monetary policy and the measure of money used in the analysis. As we shall show later in this section, these two things are intimately connected. This connection is hardly surprising since the plausibility of any given set of identifying assumptions clearly depends on the measure of money used in the statistical analysis.

To clarify the nature of the identifying assumptions which have been used in the literature, suppose that the economy evolves according to

$$(1) \quad AZ_t = B(L)Z_{t-1} + \mu_t.$$

Here Z_t denotes the time t values of the variables summarizing the state of the economic system. For now we suppose that Z_t can be partitioned as $Z_t = [Z_{1t}\ Z_{2t}]'$, where Z_{1t} denotes the time t values of the observable, endogenous nonpolicy variables in the system and Z_{2t} denotes the time t values of the policy instruments.

The fundamental sources of uncertainty in this economy are summarized by the i.i.d. random variable μ_t, which has the property that

$$(2) \quad E\mu_t\mu_t' = I,$$

where I denotes the identity matrix. The vector μ_t is partitioned as $\mu_t = [\mu_{1t}\ \mu_{2t}]'$ where μ_{it} denotes the impulses to Z_{it}, for i = 1, 2. With this notation, μ_{2t} represents the fundamental disturbances to policy. The constant matrix A summarizes the manner in which the contemporaneous values of Z_t are related to each other, while B(L) is a matrix polynomial in positive powers of the lag operator L. Notice that, absent restrictions on A or B(L), specification (1) embodies the notion that the reaction function of the policy maker, that is, the law of motion for Z_{2t}, depends on the current and past values of all the endogenous nonpolicy variables, Z_{1t}.

Now suppose we are interested in examining the historical effects of policy disturbances; that is, we want to characterize the dynamic effects of past variations in Z_{2t} arising from different values of μ_{2t}, on Z_{1t}. Given values for A and B(L), these responses can be calculated from the moving average representation of the system:

$$(3) \quad Z_t = C(L)\mu_t = \Sigma_0^{\infty} C_s \mu_{t-s},$$

where

$$(4) \quad C(L) = A^{-1}[I-B(L)]^{-1}.$$

Under our assumptions, the (k,j) element of C_s gives the response of the kth element of Z_{t+s} to a unit disturbance in the jth element of μ_t.

In practice, the problem with this procedure is that we cannot directly observe or estimate the vector of policy disturbances, μ_{2t}. The vector autoregressive representation of Z_t implied by (1) is given by

$$(5) \quad Z_t = \overline{B}(L)Z_{t-1} + v_t,$$

where

$$(6) \quad \overline{B}(L) = A^{-1}B(L),$$

$$(7) \quad v_t = A^{-1}\mu_t,$$

and

$$(8) \quad Ev_tv_t' = A^{-1}(A^{-1})' = D.$$

Absent additional restrictions on the system, all that the econometrician can hope to estimate is the parameters of $\overline{B}(L)$ and D, while the parameters A and B(L) of the moving average representation (3) are not identified. One can calculate the moving average representation implied by (5),

(9) $Z_t = \overline{C}(L)v_t$,

where

(10) $\overline{C}(L) = [I - \overline{B}(L)]^{-1}$.

However, absent very special assumptions regarding the matrix A, the statistical innovations to Z_t, namely, the v_t's, will not be the same as the fundamental disturbances to agents' environments as represented by the μ_t's. It follows that the dynamic response of nonpolicy variables in Z_t to shocks in μ_t will not coincide with the the dynamic response of those variables to shocks in v_t.[11]

In order to resolve this problem, sufficiently strong restrictions must be imposed to identify the matrix A. While a variety of procedures have been adopted by empirical analysts, the type of restrictions most relevant for the existing liquidity literature is restrictions on the contemporaneous nature of feedback between the elements of Z_t, that is, restrictions on the matrix A. To this end, most researchers in the area have proceeded by adopting a particular Wold causal interpretation of the data.[12] The general idea here is to assume that the matrix A is triangular when the variables Z_t are ordered according to their causal priority. Under this assumption, there is a unique A which satisfies (8) for a given

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[11] For a further discussion of the problems of identifying a moving average representation like (3), see Hansen and Sargent (1991).

[12] Some important exceptions are Gali (1991), King and Watson (1991), and Sims (1986,1991) who impose exclusion restrictions on the contemporaneous component of determinants of money supply and demand to identify monetary policy shocks. The first two of these also impose restrictions on the long–run relationships between the variables in their vector autoregressions. Although we do not consider the kind of identification schemes considered by these authors, they reach the same conclusion we do: that the balance of the evidence favors the conventional view that short–term interest rates fall in response to an unanticipated monetary tightening.

covariance matrix D. In the context of the liquidity literature, these types of identification schemes amount to a joint hypothesis about the nature of the contemporaneous portion of the monetary authority's feedback rule for its policy instruments and the sources of disturbances to the elements of Z_t. The set of M– and R–rules which we consider in this paper fall within this class of identification schemes.

As an example, consider the sources of identification implicit in the work of Gordon and Leeper (1991), who analyze aggregate time series data on the growth rate of the monetary base, $\Delta M0_t$; interest rates, R_t; consumer prices, P_t; and industrial production, Y_t; that is, $Z_t = [\Delta M0_t, R_t, P_t, Y_t]$. In looking for evidence of liquidity effects, Gordon and Leeper base the bulk of their inference on the moving average representation of Z_t corresponding to a lower triangular specification of the matrix A. In so doing, they identify a standardized version of the statistical innovation of $\Delta M0$ (that is, the first element in v_t) with monetary policy disturbances (that is, the first element in μ_t.) As a result, they assume monetary policy is an element of the class of M–rules, which we discussed in the introduction. The economic content of placing M0 first in the Wold causal chain is twofold: innovations to the monetary base are attributed solely to the actions of the FOMC, and in setting the growth rate of money, the FOMC does not consider the current period values of interest rates, real output, or the price level.[13] While somewhat controversial when stated in this manner, this is perfectly consistent with the long tradition of identifying the innovation in some monetary aggregate with shocks to monetary policy. (See, for example, Mishkin 1982, Barro 1981 and R. King 1991.)[14]

In sharp contrast, Bernanke and Blinder (1990) and Sims (1986, 1991) adopt a very different set of identifying restrictions which associates innovations to the federal funds

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[13] Placing R second in the Wold causal chain amounts to the assumption that time t movements in interest rate are independent of contemporaneous movements in both output and prices. Placing Y third in the Wold casual chain amounts to the assumption that time t movements in output are independent of contemporaneous movements in the price level.

[14] As Gordon and Leeper (1991) point out, there are important differences in this literature regarding which variables are allowed to enter the law of motion for the monetary aggregate.

rate with unanticipated changes in monetary policy. (They analyze elements in the class of R–rules, which we referred to in the introduction.) Working with high–order monetary aggregates, these authors adopt Wold casual interpretations of the data in which some measure of short–term nominal interest rates is placed first in the ordering.[15] The economic content of this assumption is also twofold: innovations to the federal funds rate reflect solely the decisions of the FOMC, and the contemporaneous component of the FOMC's feedback rule for setting R_t does not include objects like output, inflation, or the stock of money.

In what follows we present evidence on the liquidity effects of unanticipated changes in monetary policy using different identification schemes in conjunction with different measures of the monetary aggregate. Here, as in section 2, the three monetary aggregates considered are NBR, M0, and M1. All of the vector autoregressions we estimated included a measure of money, the federal funds rate, a measure of aggregate real output (Y), and the price level as measured by the GNP deflator.[16] Quarterly vector autoregressions including either NBR or M0 included five lags of all variables, while those including M1 included nine lags of all variables.[17]

We begin by reporting results obtained using quarterly data over the period 1959:1–1990:1. The solid lines in Figure 10 present the dynamic response of the federal funds rate to a shock in monetary policy, under five different identification schemes. The dashed lines denote two standard deviation bands about point estimates of the dynamic response functions.[18] All of the identification schemes share the assumption that the

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[15] Sims (1986, 1991) works exclusively with M1 while Bernanke and Blinder (1990) experiment with both M1 and M2.

[16] We found that our results were not affected when we used either the consumer price index or the constant weighted GNP deflator instead of the GNP deflator as our measure of the price level.

[17] Lag lengths were selected based on evidence regarding the serial correlation in the error term in the vector autoregression, as measured using the Q statistic discussed in Doan (1990).

[18] These were computed using the method described in Doan (1990), example 10.1, using 100 draws from the estimated asymptotic distribution of the vector autoregressive coefficients.

unanticipated change in monetary policy is some orthogonalized component of the innovation to the monetary aggregate included in the vector autoregression. The three columns contain results pertaining to the case in which the measure of money is NBR, M0, and M1.

Each of the five rows displays the dynamic response of the federal funds rate to an unanticipated change in monetary policy generated under a different identification scheme, each of which is summarized by the label "RESP of FF to M/X", M = {NBR, M0, M1} and X = {0, R, Y, P, (P,Y)}. The Wold ordering underlying the first row corresponds to {M, R, Y, P}. Placing M first in the ordering equates, after scaling by their standard deviation, innovations in M to unanticipated changes in monetary policy. This corresponds to the assumption that the contemporaneous portion of the monetary authority's feedback rule for setting M_t does not involve R_t, Y_t or P_t. The Wold ordering underlying the second row corresponds to {R, M, Y, P}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in M_t which is orthogonal to the innovation in R_t. This corresponds to the assumption that the contemporaneous portion of the monetary authority's feedback rule involves R_t, but not P_t or Y_t. The Wold ordering underlying the third row is given by {Y, M, R, P}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in M_t which is orthogonal to the innovation in Y_t, that is, the contemporaneous portion of the monetary authority's feedback rule for M_t involves Y_t, but not R_t or P_t. The Wold ordering underlying the fourth row is given by {P, M, R, Y}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in M_t which is orthogonal to the innovation in P_t, that is, the contemporaneous portion of the monetary authority's feedback rule for M_t involves P_t, but not R_t or Y_t. Finally, the Wold ordering underlying the fifth row is {P, Y, M, R}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in M_t which is orthogonal to innovations in P_t and Y_t; that is, in setting M_t the monetary authority looks at P_t and Y_t, but not R_t. In no case do we impose any restrictions on the lagged components of the vector autoregression (VAR).

Consider first our results with NBR. Notice that, regardless of which identification scheme is imposed, innovations to monetary policy are always followed by sharp, persistent, statistically significant declines in R_t. In all but one case, the dynamic response is the same: the immediate impact of the shock to monetary policy is to drive down R_t, which stays below its preshock level for approximately 16 quarters. In the one exception, the second row, labeled "RESP of FF to NBR/R," the identification scheme rules out, a priori, a contemporaneous response of R_t. Even here, though, R_t falls in the period after the shock and stays below its preshock level for approximately 20 quarters.

The second and third columns of Figure 10 reveal that changing the measure of money from NBR to either M0 or M1 has a drastic impact on inference. Measured this way, innovations to monetary policy are followed by increases in R_t. In the case of M0, the basic patterns are very similar across the different identification schemes: the immediate impact of the shock is to drive up R_t, which stays above its preshock level for approximately eight quarters, but then falls and stabilizes at a level slightly below its preshock level.[19] Notice, though, that the standard errors associated with these impulse response functions are quite large. Indeed, one cannot reject, at reasonable confidence levels, the null hypotheses that the federal funds rate rises, falls, or is unchanged following a shock in M0. Evidently, this monetary aggregate contains very little information for the federal funds rate. When M1 is used as the measure of money, the positive response of R_t to a policy shock is statistically more significant than is the case for M0.

In order to assess the sensitivity of results to the use of quarterly data, we redid our analysis using monthly data. These results are presented in Figure 11. Since real GNP data are unavailable at the monthly level, we used industrial production as our measure of aggregate output and the consumer price index as our measure of the aggregate price level.

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[19] The identification scheme labeled "RESP of FF to M0/R" precludes an immediate reaction of R_t to a shock in monetary policy.

As can be seen, these changes have very little impact on our results. Orthogonalized shocks to NBR continue to drive the federal funds rate down, M0 continues to contain very little information for the federal funds rate, and orthogonalized shocks to M1 drive the federal funds rate up.

To address the issue of sample period sensitivity we also redid our analysis using two distinct sample periods. Figures 12 and 13 report results for NBR using quarterly and monthly data, respectively. Figures 14 and 15 report the analogous results for M0 while Figures 16 and 17 display the analogous results for M1. The three columns of each figure display the response of R_t to unanticipated shocks to monetary policy, for the periods 1959:1–1990:1, 1959:1–1978:4, and 1979:1–1990:1, respectively. The class of identification schemes considered is the same as that underlying Figures 10 and 11.

Consider first our results for NBR. Figures 12 and 13 reveal that there is some sample period sensitivity. For both the monthly and quarterly data, the pre–1979 and post–1979 results appear quite different from the full sample results. Still, for the monthly data, it remains true that R_t always drops following an expansionary monetary disturbance. This is true regardless of which identification scheme or which sample period is adopted. At the same time, the persistence of the drop in the federal funds rate seems much shorter in the post–1979 period. Figures 14 and 15 reveal that the response functions generated using M0 seem to be relatively stable across sample periods. Figures 16 and 17 reveal that the results obtained using M1 are the most sensitive to splitting the sample. Here, as with NBR, the pre–1979 period looks similar to the entire sample period, while the post–1979 period looks quite different.

To summarize this portion of our analysis, we find that when NBR is used as the measure of money, there is very strong evidence that, relative to the identification schemes considered, unanticipated expansionary changes in monetary policy drive the federal funds rate down. It is hard to imagine reconciling this finding with models that do not incorporate liquidity effects. In sharp contrast, when either M0 or M1 is used as the measure of money, unanticipated expansionary changes in monetary policy drive the federal funds rate up. Evidently, a given class of identification schemes generates very different results for different measures of money. This result is hardly surprising. As we stressed in the introduction, there is no reason to believe that a given set of identifying assumptions will be equally appropriate across high– and low–order measures of the monetary aggregate. Aggregates like NBR, M0, and M1 are influenced by very different sets of economics agents.

Given these apparently conflicting results, how can we realistically hope to proceed? One response is to investigate whether certain combinations of identification assumptions and monetary aggregates can be eliminated on the basis of their implications for variables other than interest rates. This seems to be the case for the class of identification schemes which equate unanticipated changes in monetary policy with some fraction of the innovations in M0 or M1. For example, Figure 18 displays the dynamic response function of GNP to shocks in monetary policy for the M–rule class of identification schemes underlying Figures 10 and 11. Notice that when M1 is used in the analysis, shocks to monetary policy are followed by sharp persistent declines in GNP which last over nine years. While less pronounced with M0, the salient effect of such shocks is also a large persistent decline in GNP.[20] Even taking sampling uncertainty into account these declines last for roughly three years, with the exact horizon depending on the identification scheme used. In sharp contrast, when NBR is used in the analysis, these types of shocks generate increases in aggregate output.[21] We conclude that, if one conditions on the class of M–rules considered here, then the results based on NBR are the most plausible. Those results provide strong support for the importance of the liquidity effect.

A different class of identifying restrictions not captured in Figures 10–17 emerges

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[20] This effect emerges most clearly under the assumption that the monetary authority sets M0 taking into account the contemporaneous values of the price level and/or GNP.

[21] There is some sensitivity to when the rise in GNP begins.

from the analyses of McCallum (1983), Bernanke and Blinder (1990), and Sims (1986, 1991). These authors argue that, at least relative to high–order monetary aggregates like M1 or M2, the innovation in R_t is a better measure of unanticipated changes in monetary policy than the innovation to the stock of money. In pursuing this idea, Bernanke and Blinder (1990) and Sims (1986, 1991) assume that innovations in R_t arise solely from the actions of the monetary authority and that the contemporaneous portion of the feedback rule for setting R_t does not include M_t, Y_t, or P_t. In short, they place R_t first in their Wold causal chain. More generally, their arguments suggest measuring monetary policy as some orthogonalized component of the innovation in the federal funds rate.

Figure 19 reports a subset of the implications of this class of identifying restrictions. In particular, each row displays the response function of R_t to a shock in monetary policy measured using a different identification scheme. The graphs labeled "RESP of M to FF/X," M = {NBR, M0, M1} and X = {0, R, Y, P, (P,Y)} denote the response of the monetary aggregate to a contractionary monetary policy disturbance, where the latter is identified as that component of the innovation in R_t which is orthogonal to X. The Wold ordering underlying the first row is {R, M, Y, P}. This corresponds to the identification scheme imposed by Bernanke and Blinder (1990) as well as Sims (1986, 1991). The Wold ordering underlying the second row is {M, R, Y, P}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in R_t which is orthogonal to the innovation in M_t. This corresponds to the assumption that the contemporaneous portion of the monetary authority's feedback rule for setting R_t involves M_t, but not P_t or Y_t. The Wold ordering underlying the third row is {Y, R, M, P}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in R_t which is orthogonal to the innovation in Y_t. This corresponds to the assumption that the only contemporaneous variable which the FOMC looks at when setting R_t is Y_t. The Wold ordering underlying the fourth row is {P, R, M, Y}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in R_t which is orthogonal to the innovation in P_t. This corresponds to the assumption that the only contemporaneous variable which the FOMC looks at when setting R_t is P_t. Finally the Wold ordering underlying the entries of the fifth row is {P, Y, R, M}, so that the unanticipated change in monetary policy is measured as that portion of the innovation in R_t which is orthogonal to the innovations in both P_t and Y_t. This corresponds to the assumption that in setting R_t, the FOMC looks at P_t and Y_t, but not M_t.

Figure 19 reveals that, with one important exception, unanticipated changes in monetary policy, corresponding to an increase in R_t, are followed by long declines in the stock of money, regardless of whether the latter is measured as NBR, M0 or M1. The only exception to this pattern arises with NBR under the identification scheme generating the graph labeled "RESP of NBR to FF/NBR." Here NBR rises for approximately 15 quarters before falling below its pre shock level. One possible interpretation of this uses the fact that the monetary shock here is that component of the innovation to R_t which is orthogonal to NBR, that is, it is the movement in R_t which cannot be predicted on the basis of the current level of NBR. Viewed from this perspective, the increase in NBR after such a shock may be the consequence of a policy in which the monetary authority smooths fluctuations in the federal funds rate arising from what it perceives to be shocks to the demand for money. Goodfriend (1991) and others argue forcefully that this has been an important feature of postwar federal feserve policy. Finally, Figure 20 presents the analog to Figure 19 obtained with monthly data. As can be seen, our results are quite insensitive to this perturbation. Thus, abstracting from one identification scheme, conditioning on the R–rules leads to the inference that there is an important liquidity effect.

4. Summary

We conclude this paper by summarizing our two main findings.

First, we found that when nonborrowed reserves are used as the measure of money, inference regarding the effects of unanticipated changes in monetary policy on interest rates is very robust. When the shock to monetary policy is measured as some orthogonalized component of the innovation to nonborrowed reserves, the federal funds rate displays a sharp, large, persistent decline in response to expansionary monetary policy. When the shock to monetary policy is measured as some orthogonalized component of the innovation to the federal funds rate, unanticipated contractionary monetary policy, corresponding to an increase in the federal funds rate, is accompanied by a sharp, large persistent decline in NBR.

To us, these findings constitute strong evidence in favor of the view that unanticipated expansionary open market operations drive interest rates down, at least in the short run, that is, the federal reserve lowers interest rates by withdrawing nonborrowed reserves from the system. It seems unlikely that these findings can be reconciled with models that do not incorporate liquidity effects.

Second, we found that when either M0 or M1 is used as the measure of money, inference regarding the effects of unanticipated changes in monetary policy on interest rates is very sensitive to the identification scheme adopted. When the shock to monetary policy is measured as some orthogonalized component of the innovation to the federal funds rate, unanticipated contractionary monetary policy, corresponding to an increase in the federal funds rate, is accompanied by a sharp, large, persistent decline in M0 and M1. Thus, this class of monetary rules generates evidence in support of liquidity effects. In contrast, when the shock to monetary policy is measured as some orthogonalized component of the innovation to M0 or M1, the federal funds rate displays a large, persistent rise in response to expansionary monetary policy. However, these shocks generate implications for real output which seem to us implausible. We conclude that the balance of the evidence, including the dynamic correlations discussed in section 2, is consistent with the conventional view of the effects of monetary policy disturbances on interest rates.

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Figure 1: Three Short Term Interest Rates, 1959:1-1990:1

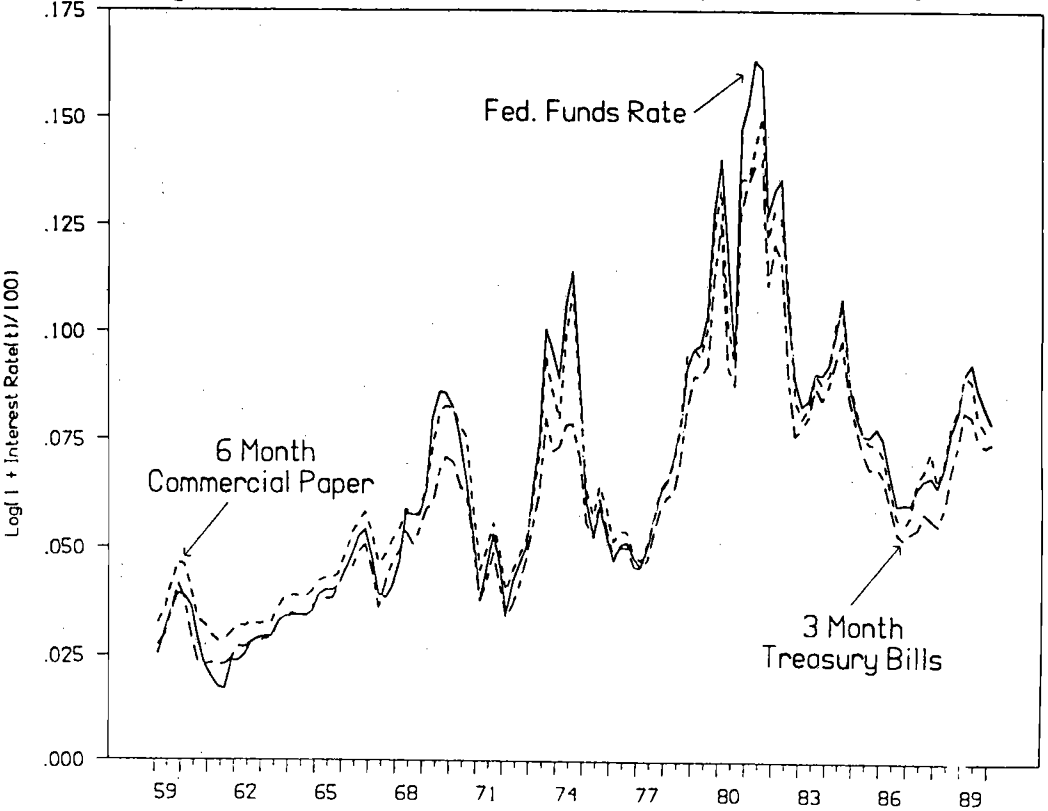


Figure 2a: Nonborrowed Reserves with HP Trend, 1959:1-1990:1

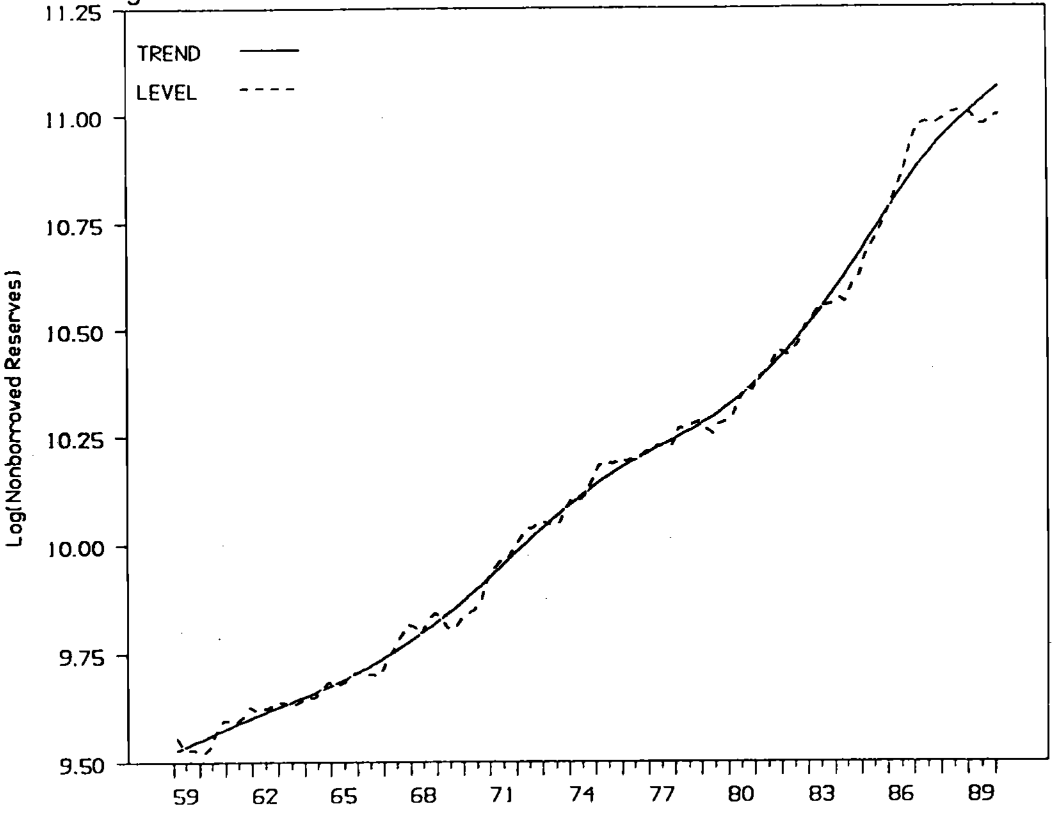
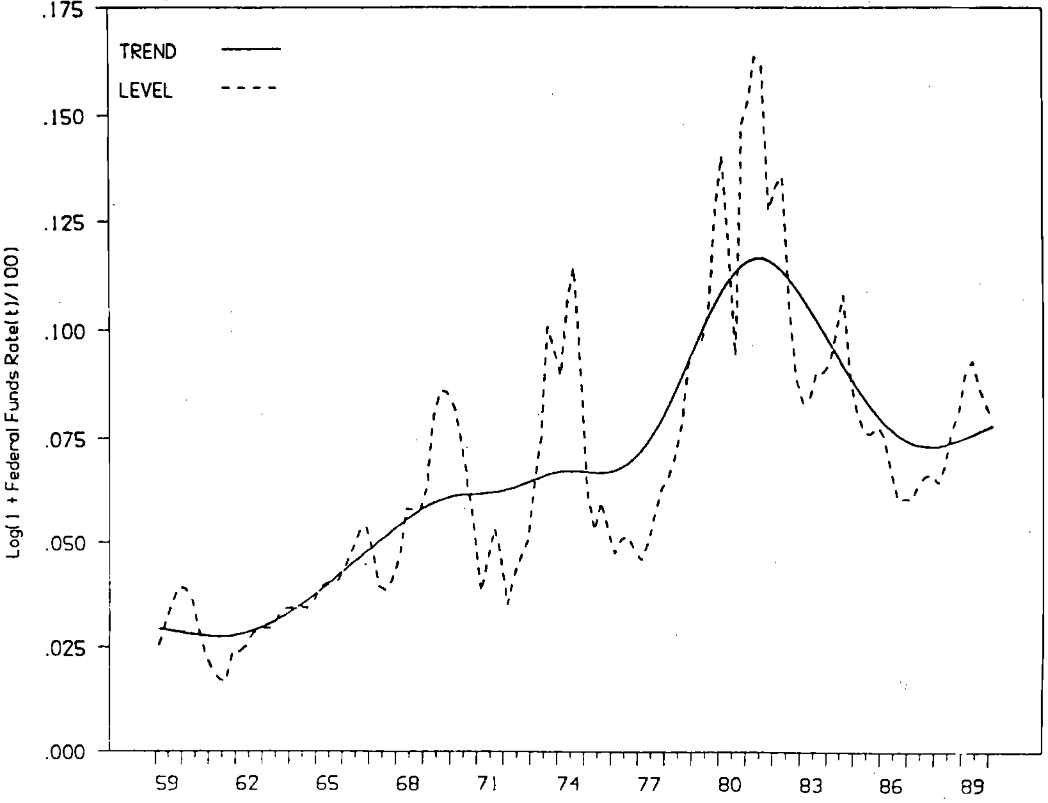


Figure 2b: Federal Funds Rate with HP Trend, 1959:1-1990:1



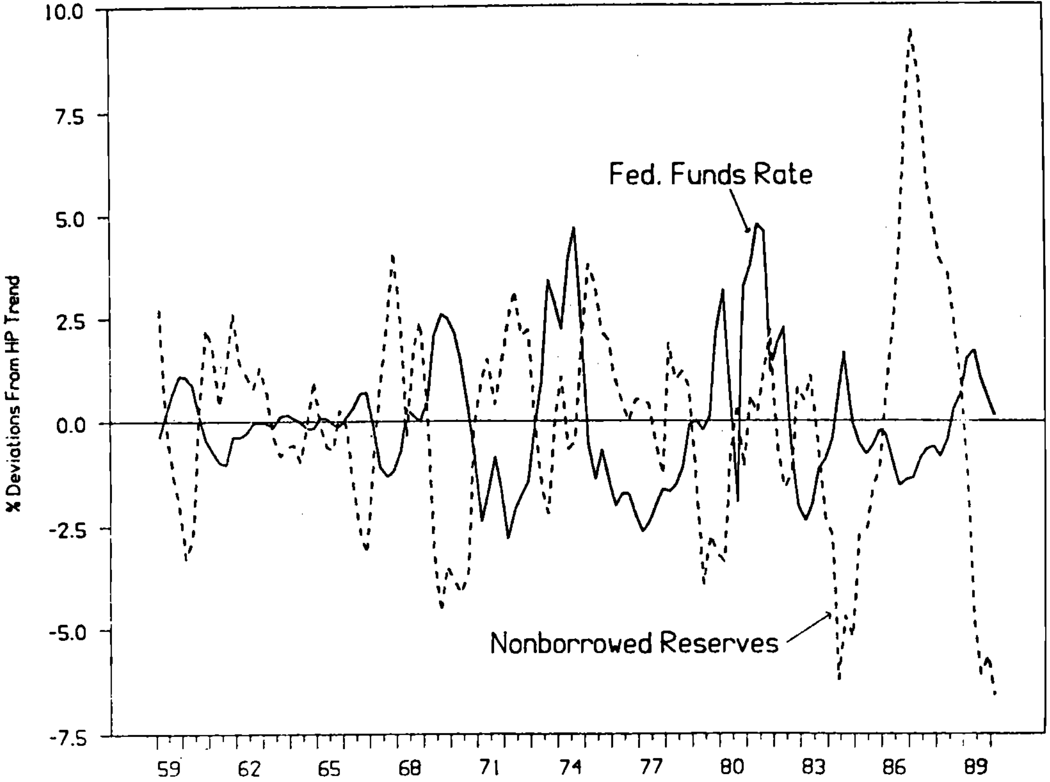


Figure 3a: HP Filtered Federal Funds Rate and Nonborrowed Reserves, 1959:1-1990:1

Figure 3b: HP Filtered Federal Funds Rate
and Real GNP, 1959:1-1990:1

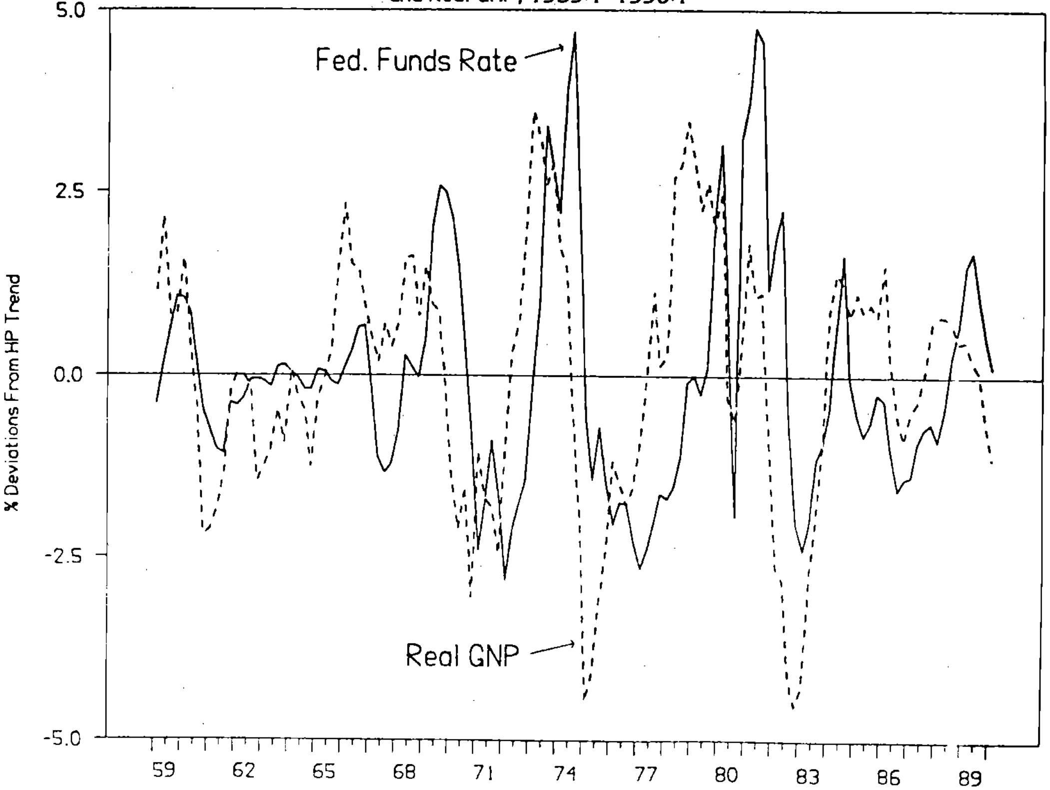
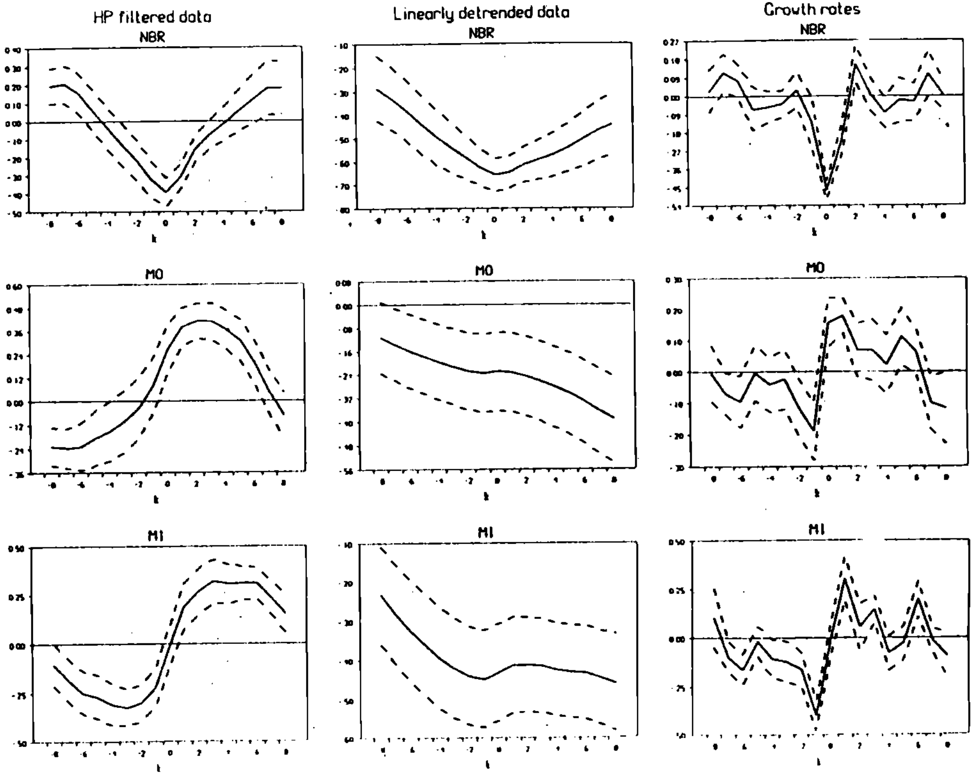


Figure 4: Correlation (Federal Funds Rate(t), M(t-k)).
k = -8,...,8, M = NBR, M0, M1, 1959:1-1990:1.


HP filtered data

Linearly detrended data

Growth rates

NBR

M0

M1

Figure 5: Correlation (Federal Funds Rate(t), M(t-k)),
k = -8,...,8, M = BR, M0 - BR, M1 - BR, 1959:1-1990:1.

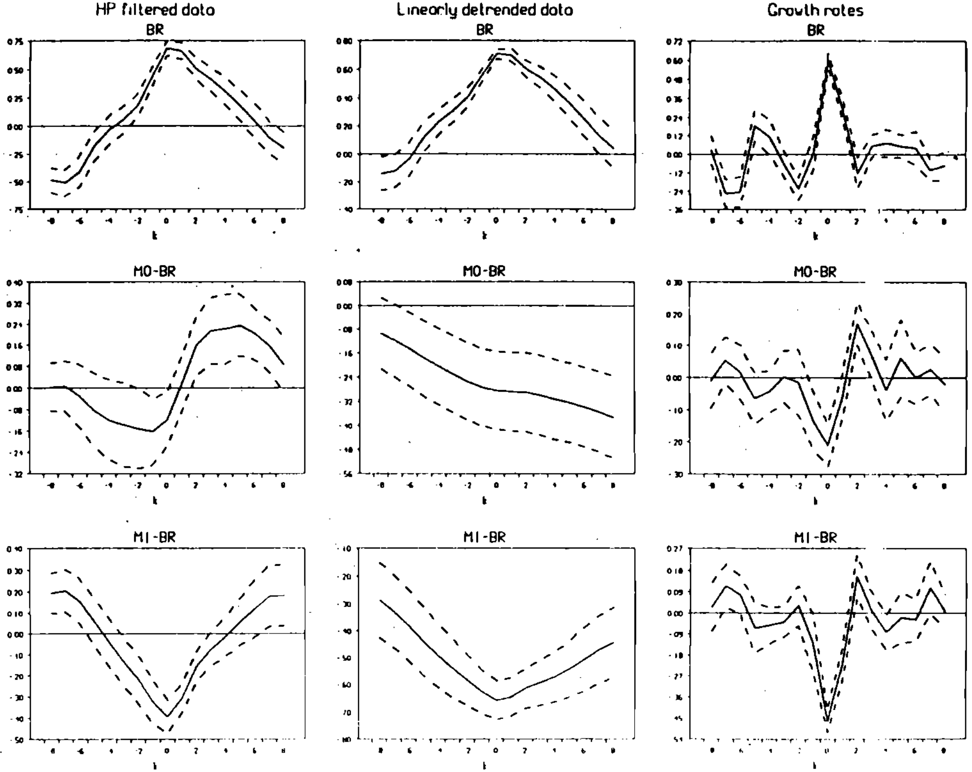
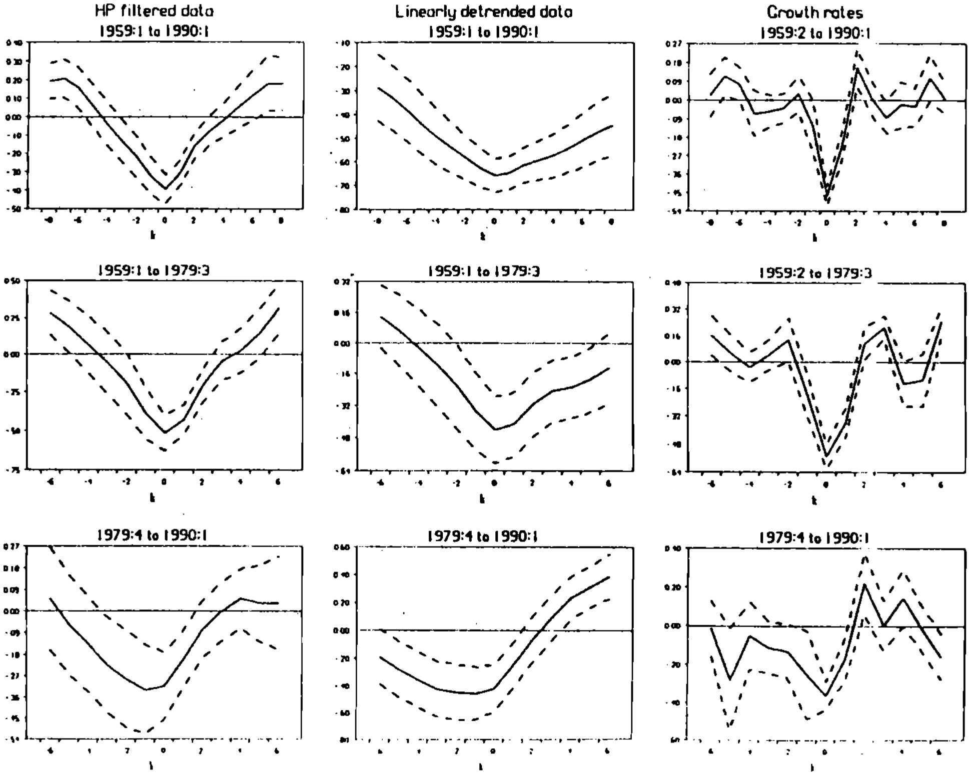


Figure 6: Correlation (Federal Funds Rate(t), NBR(t-k)),
k = -8,...,8.



HP filtered data 1959:1 to 1990:1

Linearly detrended data 1959:1 to 1990:1

Growth rates 1959:2 to 1990:1

1959:1 to 1979:3

1959:1 to 1979:3

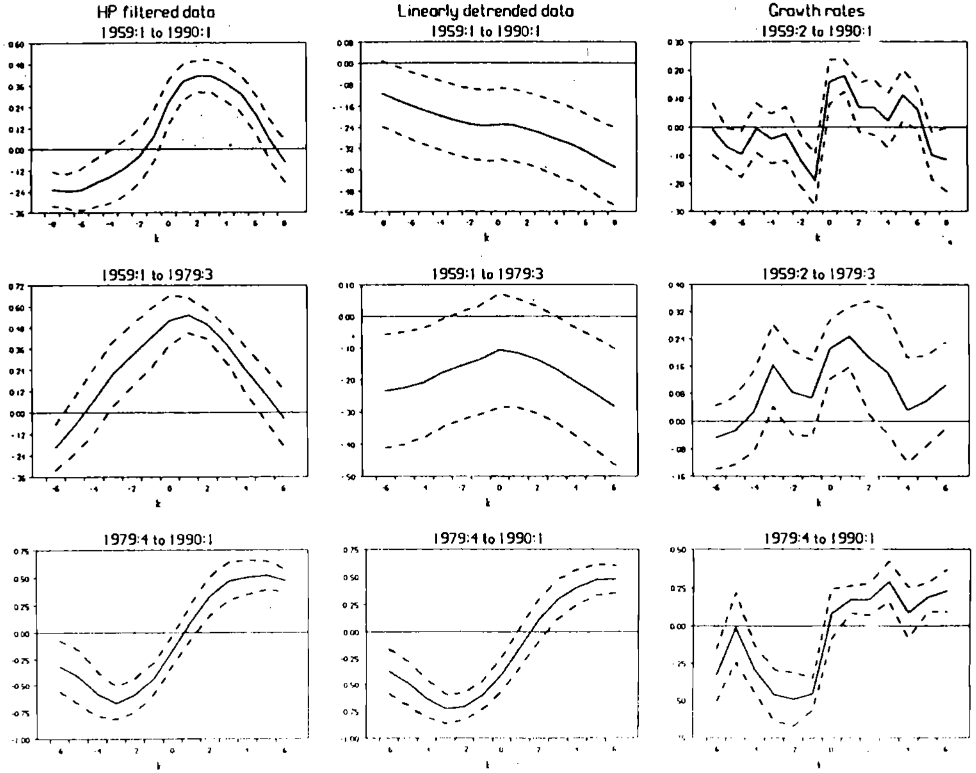
1959:2 to 1979:3

1979:4 to 1990:1

1979:4 to 1990:1

1979:4 to 1990:1

Figure 7: Correlation (Federal Funds Rate(t), M0(t-k)), k = -8,...,8.



HP filtered data 1959:1 to 1990:1

Linearly detrended data 1959:1 to 1990:1

Growth rates 1959:2 to 1990:1

1959:1 to 1979:3

1959:1 to 1979:3

1959:2 to 1979:3

1979:4 to 1990:1

1979:4 to 1990:1

1979:4 to 1990:1

Figure 8: Correlation (Federal Funds Rate(t), M1(t-k)), k = -8,...,8.

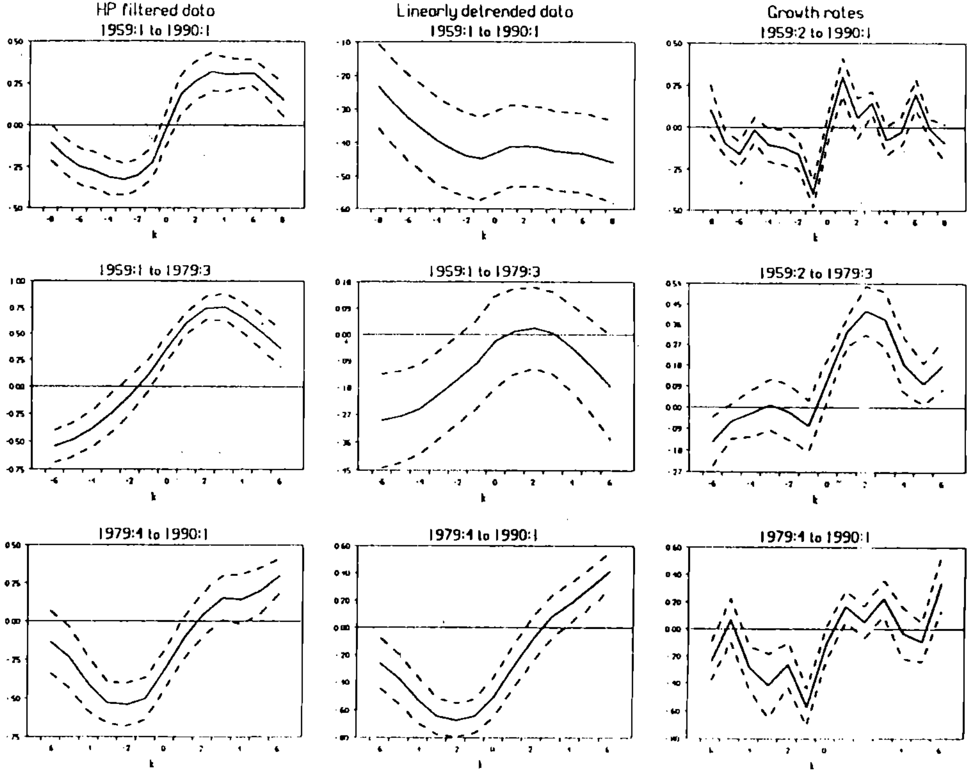


Figure 9: Correlation (Federal Funds Rate(t), GNP(t-k)).
k = -8,...,8.

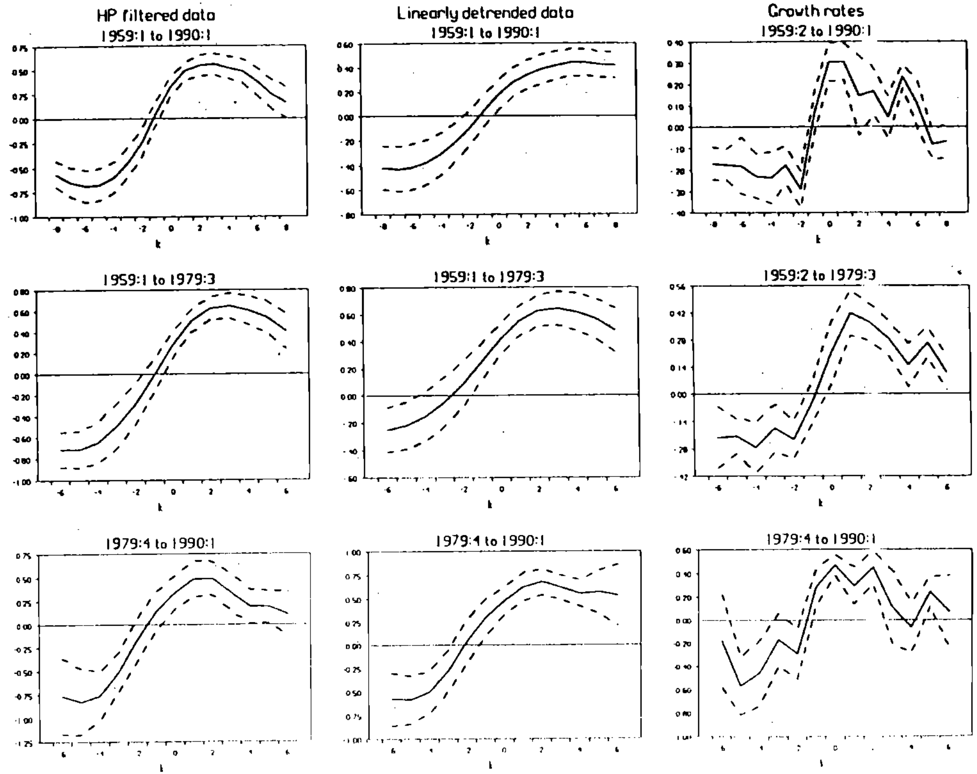


Figure 10: Response of Federal Funds Rate to Policy Shock
Under M - Rule (Various Measures of Money, Quarterly, 1959:1-1990:1)

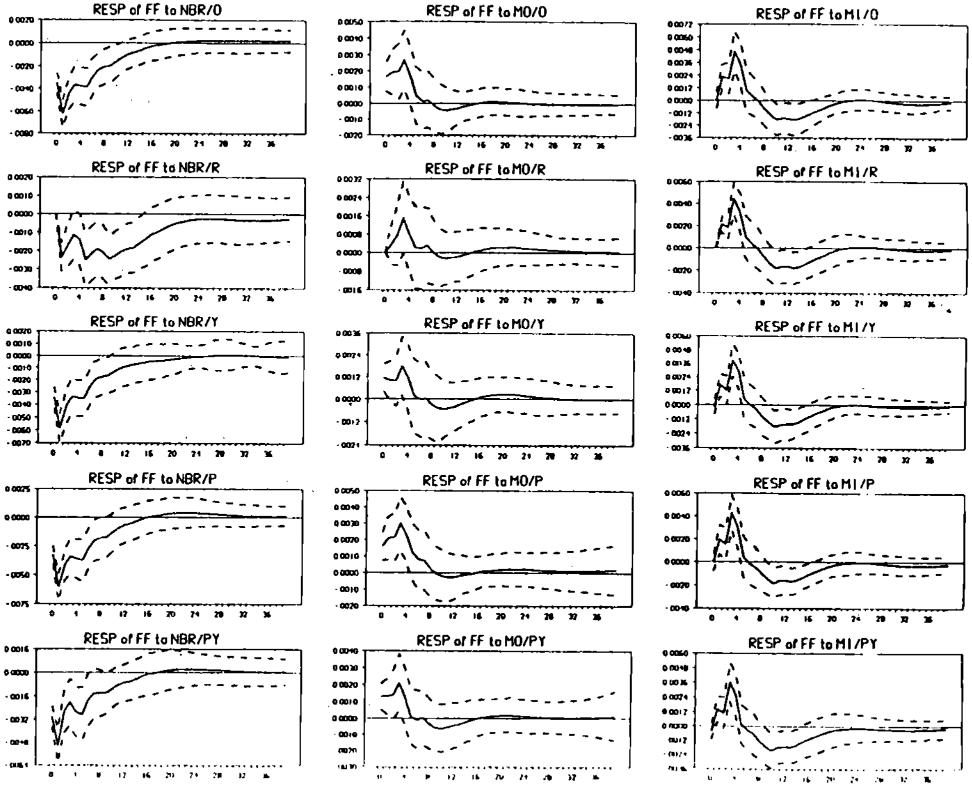


Figure 11: Response of Federal Funds Rate to Policy Shock
Under M - Rule (Various Measures of Money, Monthly, 1959:1-1990:1)

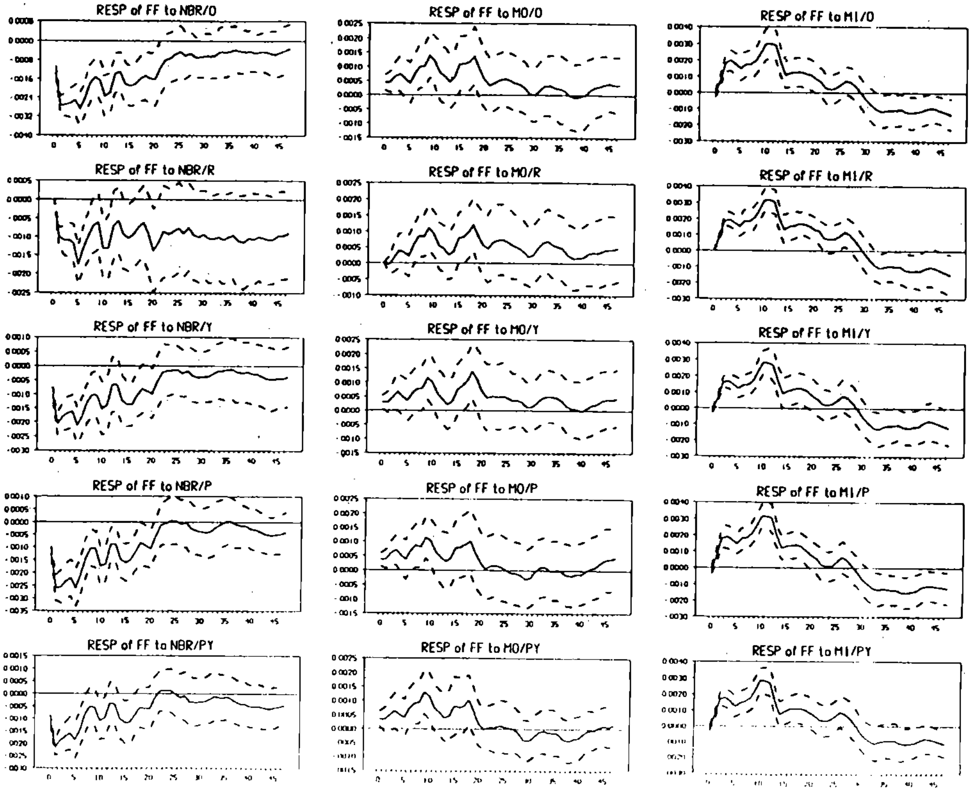


Figure 12: Response of Federal Funds Rate to Policy Shock
Under M - Rule (NBR, Quarterly, Various Samples)

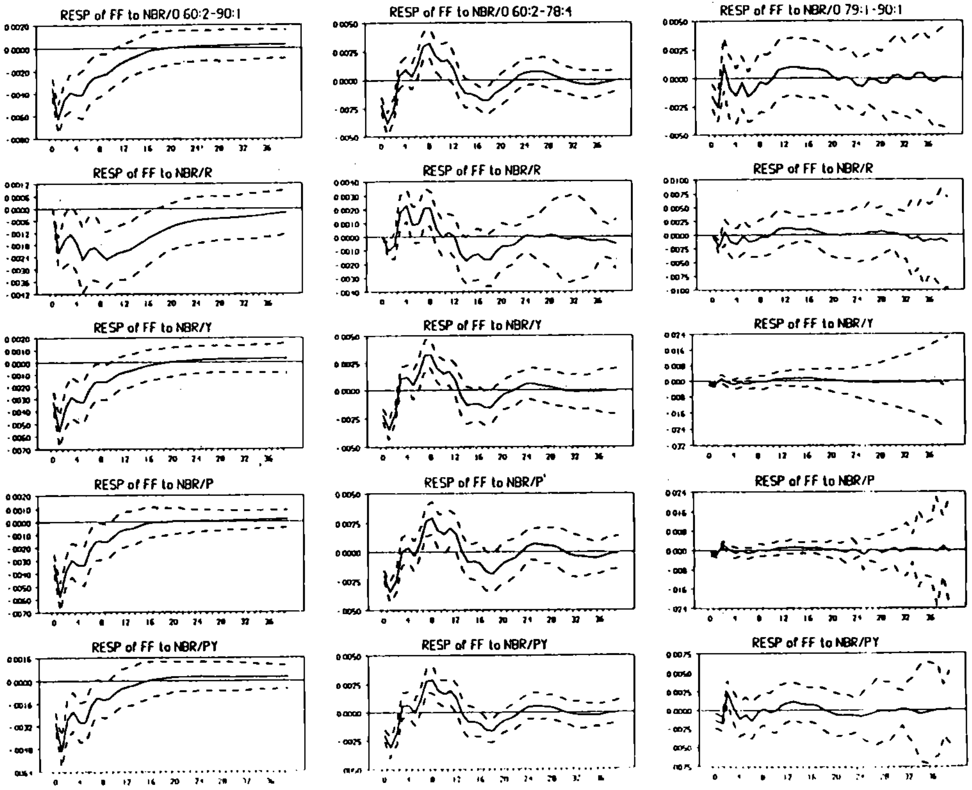


Figure 13: Response of Federal Funds Rate to Policy Shock
Under M - Rule (NBR, Monthly, Various Samples)

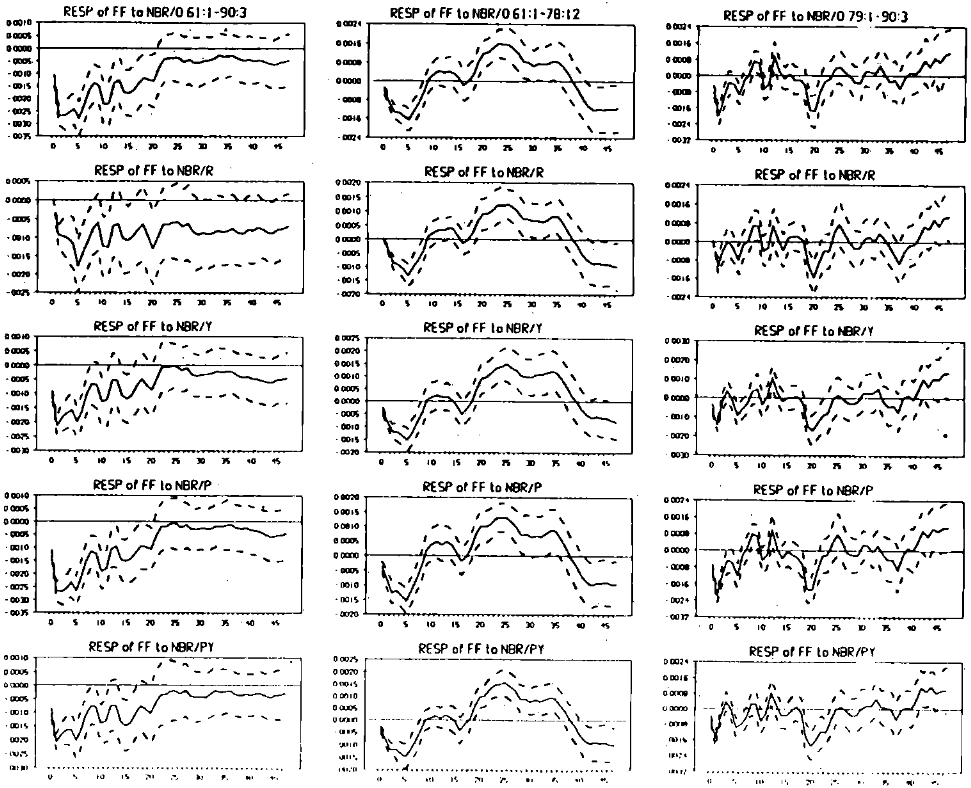


Figure 14: Response of Federal Funds Rate to Policy Shock
Under M - Rule (M0, Quarterly, Various Samples)

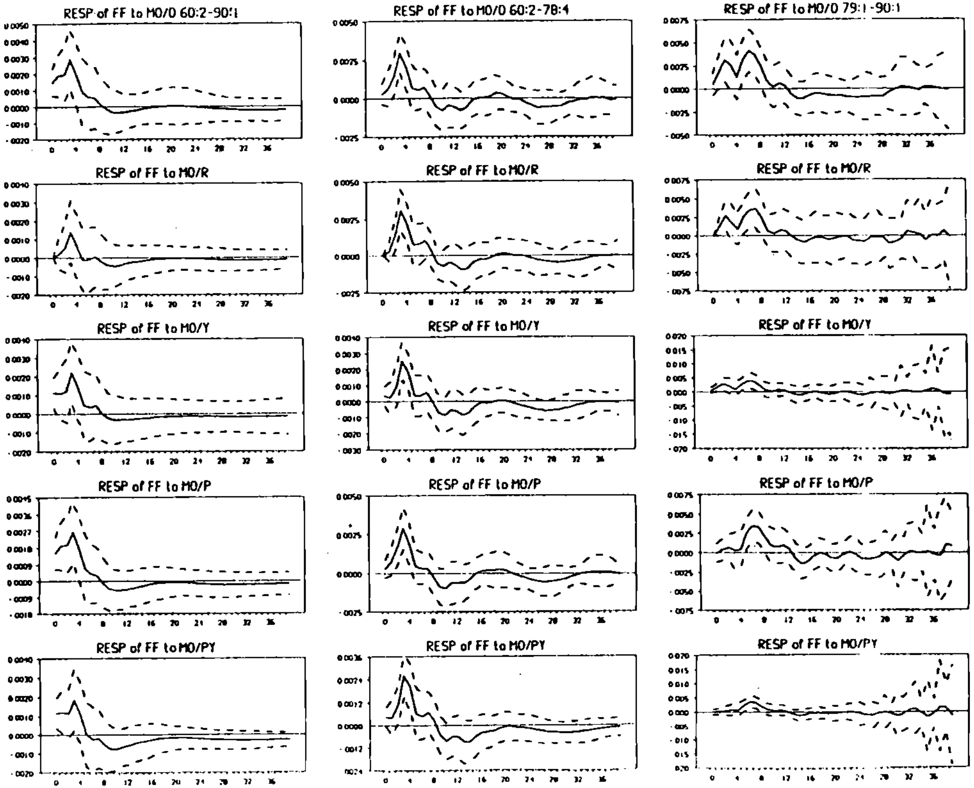


Figure 15: Response of Federal Funds Rate to Policy Shock
Under M - Rule (M0, Monthly, Various Samples)

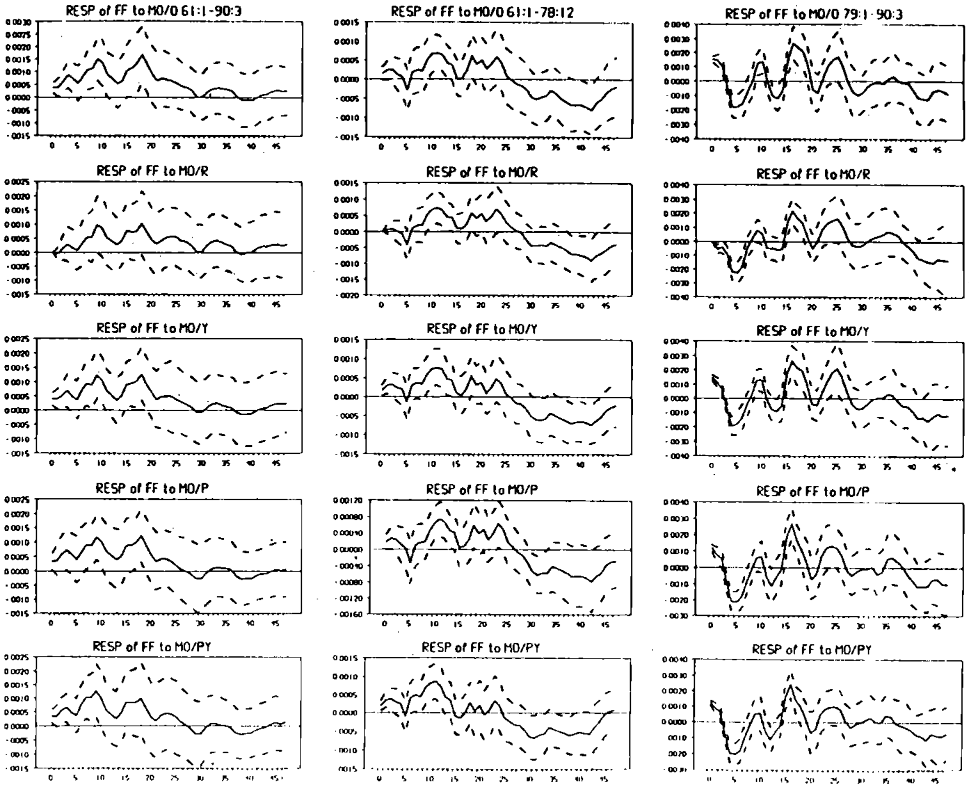


Figure 16: Response of Federal Funds Rate to Policy Shock
Under M - Rule (M1, Quarterly, Various Samples)

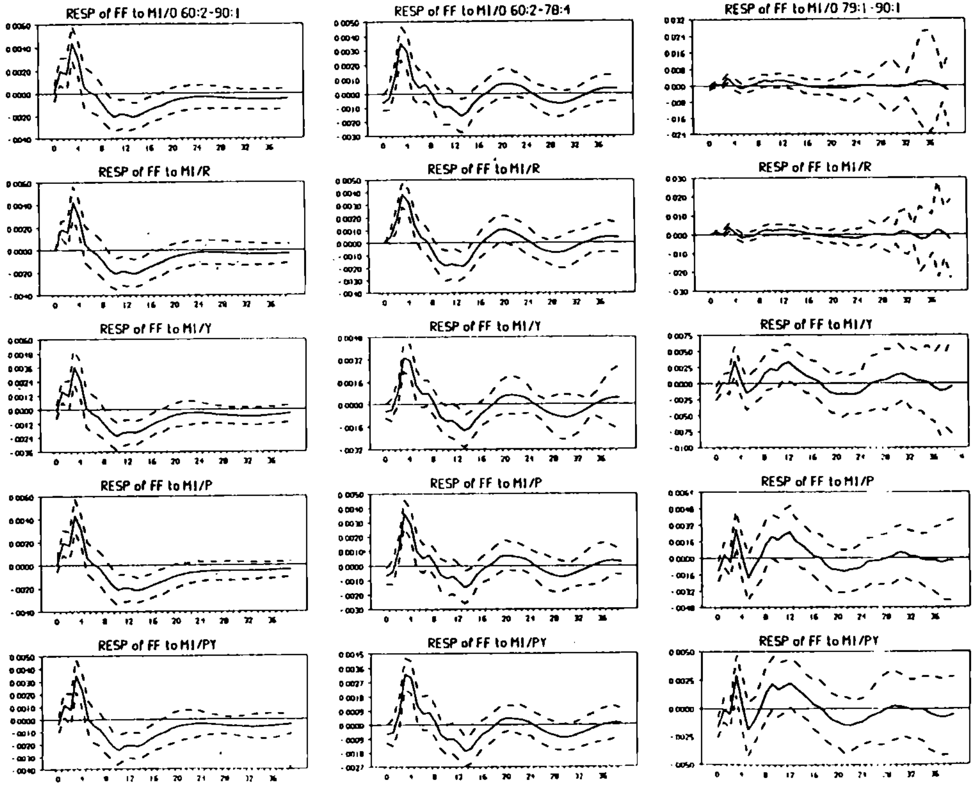


Figure 17: Response of Federal Funds Rate to Policy Shock Under M - Rule (M1, Monthly, Various Samples)

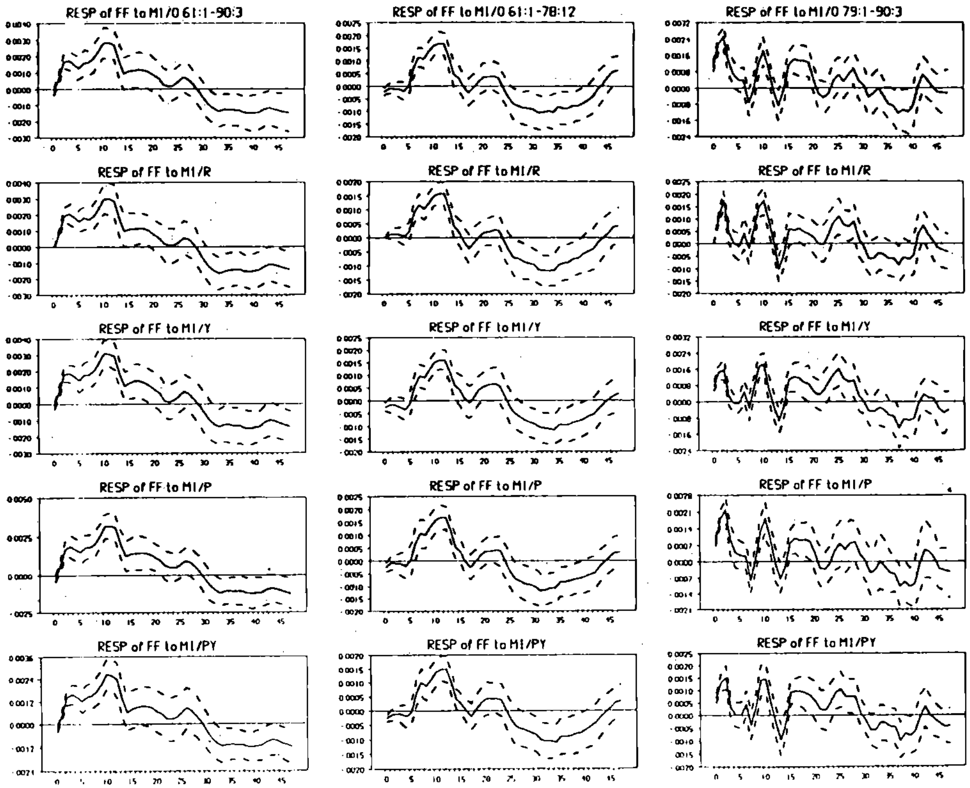


Figure 18: Response of GNP to Policy Shock Under M - Rule
(Various Measures of Money, Quarterly, 1959:1 - 1990:1)

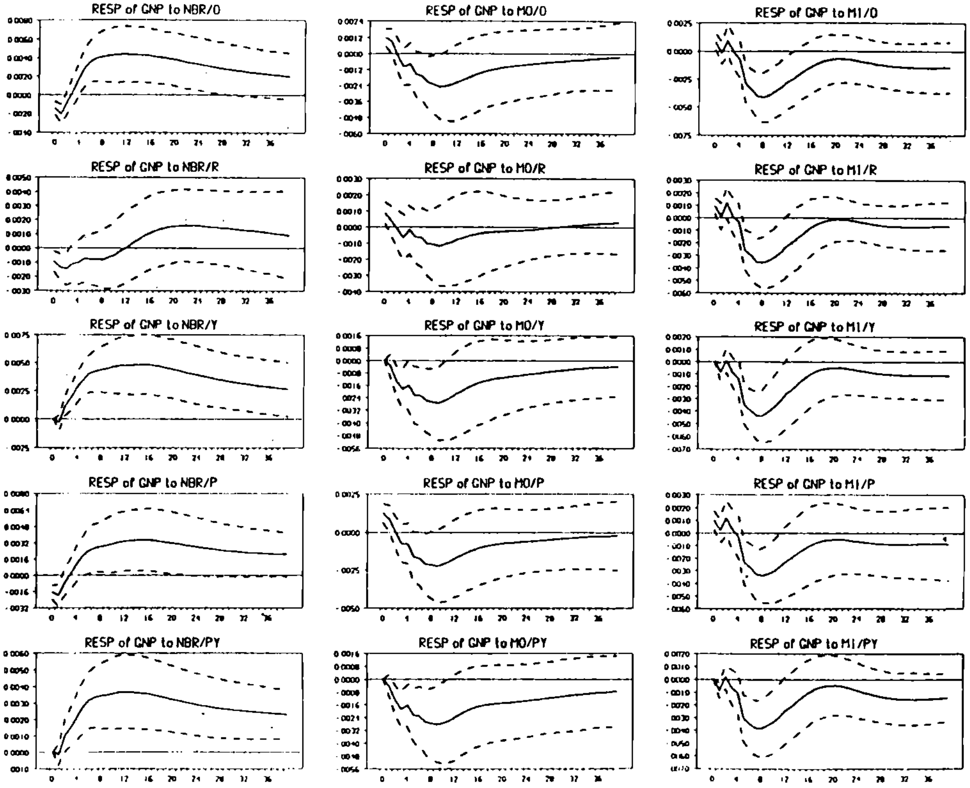


Figure 19: Response of Money to Policy Shock Under R - Rule
(Various Measures of Money, Quarterly, 1959:1 - 1990:1)

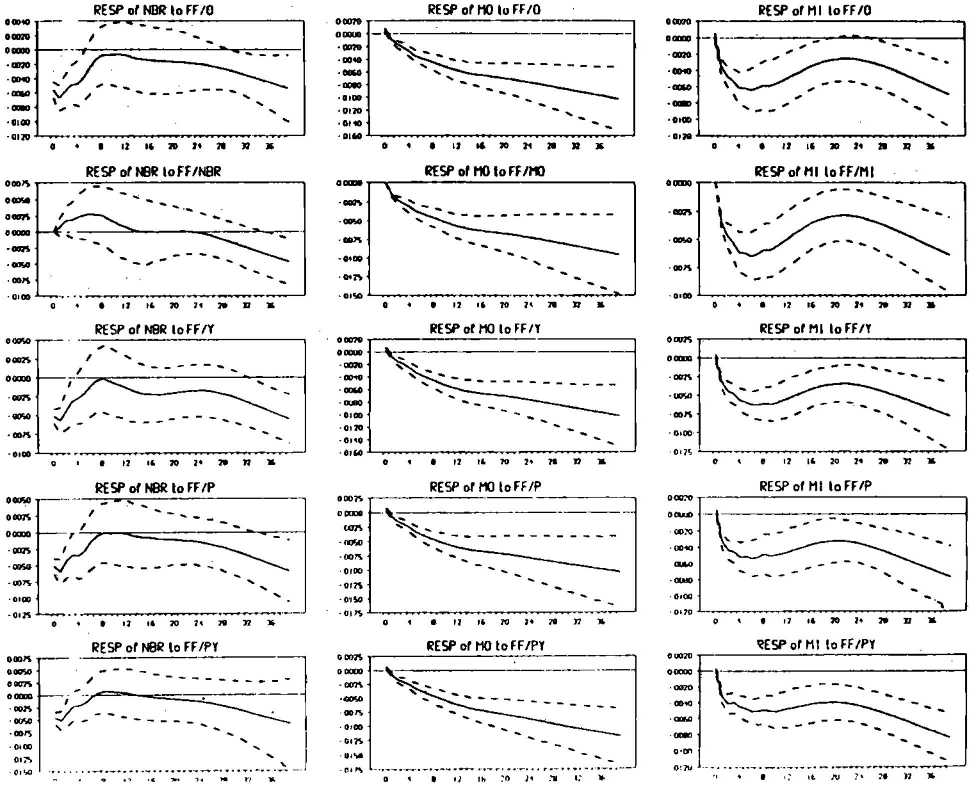


Figure 20: Response of Money to Policy Shock Under R - Rule
(Various Measures of Money, Monthly, 1959:1 - 1990:1)