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Estimating the effects of monetary shocks: An evaluation of different approaches

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Abstract

This paper compares several methods for estimating the effects of monetary innovations on key macroeconomic variables and, subsequently, clarifies issues related to the use of instrumental variables in the identification of structural impulse responses. In particular, we make explicit the property that a measure of monetary policy must satisfy in order to identify the effects of monetary shocks. Within our framework we find that none of the currently popular methods of identifying the effects of monetary shocks are supported by the data. We also indicate how current approaches can be combined to provide unbiased estimates of the effects of monetary disturbances. © 1998 Elsevier Science B.V. All rights reserved.

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1. Introduction

Quantifying the effects of monetary shocks is one of the most important questions of empirical monetary economics. Currently there are several competing approaches in vogue. On the one hand, there are schemes based on the Choleski decomposition of VAR residuals. Within this approach, the main

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debated issue relates to whether there exists an indicator of monetary policy which can justifiably be placed within a Wold-causal ordering. For example, in his seminal work on VAR systems, Sims (1980) uses M1 as his indicator of monetary policy. More recently, Strongin (1995) and Christiano and Eichenbaum (1992) have argued in favor of the use of non-borrowed reserves, while Bernanke and Blinder (1992) and Sims (1992) suggest the use of the Federal Funds rate. On the other hand, the approach favored by Romer and Romer (1989, 1990) is to use episodes viewed as monetary contractions as a means of identifying monetary shocks. Here the main issue of contention is whether such episodes can reasonably be considered exogenous with respect to other shocks to the economy.

In this paper we present a simple framework in which these competing approaches can be evaluated. We emphasize that the debate regarding the appropriate choice of a monetary policy variable is intimately linked with the possibility of using variables, such as the Romer dummies, as a means of identifying the monetary business cycle. In effect, we show that access to a variable that is correlated with monetary shocks and uncorrelated with any other shocks does not necessarily allow one to identify the effects of monetary innovations using any arbitrary measure of monetary policy. In particular, our framework makes explicit the property that a measure of monetary policy must satisfy to allow the identification of the monetary business cycle.

Our empirical strategy consists in exploiting the episodes counted as monetary contractions by Romer and Romer in order to examine the issue within a unified instrumental variable framework. In contrast to Romer and Romer (1989, 1990), we focus on identifying the structural impulse responses associated with a monetary shock and on testing the underlying restrictions. The instrumental variable approach we adopt allows us to examine whether any of the contested Choleski decompositions is appropriate to identify the monetary business cycle. The approach also allows us to estimate (under certain conditions) the effect of monetary shocks even when Choleski decompositions are inappropriate. Since it is not a priori clear that the Romer dummy variables are appropriate instruments for the identification of monetary shocks, we take special care to provide new evidence of their validity. In particular, our approach differs from previous attempts to use these dummy variables as instruments in that we make explicit the conditions under which our strategy is appropriate and we report test statistics associated with these conditions.

The main findings of the paper are that (i) the currently popular methods of identifying the effects of monetary shocks are not supported by the data, and (ii) unbiased estimates of the effects of monetary shocks can be estimated by a procedure that includes instrumenting non-borrowed reserves with the dummy variables constructed by Romer and Romer (1989), and (iii) the failure to take into account the endogeneity of monetary policy variables, including non-borrowed reserves, has probably led to under-estimates of the liquidity effect.

The paper is structured as follows. In Section 2 we present a statistical framework which helps clarify certain issues related to the identification of structural impulse responses. In Section 3 this framework is used to compare different approaches currently used to identify the monetary business cycle. Section 4 offers concluding comments.

2. Econometric framework

This section begins by making explicit conditions under which instrumental variables can be used to identify structural impulse responses. We then indicate how the proposed framework can be used to test whether a Choleski decomposition is likely to be preferable to an instrumental variable approach. The analysis highlights why the choice of an appropriate indicator of monetary policy remains a crucial issue even if one has access to seemingly valid instruments.

Consider the statistical model comprised of n stationary variables represented by the $n \times 1$ vector $y(t)$ and n fundamental shocks $\varepsilon(t)$, where $E\varepsilon(t) = 0$, $E[\varepsilon(t)\varepsilon(t)']$ is a diagonal matrix and $E[\varepsilon(t)\varepsilon(t-j)'] = 0$ for $j \neq 0$. The structural relationship between the variables $y(t)$ and the fundamental shocks $\varepsilon(t)$ is assumed to be captured by the following representation.

$$y(t) = Ay(t) + B(L)y(t-1) + \Pi\varepsilon(t). \tag{2.1}$$

In Eq. (2.1), A is an $n \times n$ matrix with zeroes on the diagonal, $B(L)$ is an $n \times n$ polynomial in the lag operator and Π is an $n \times n$ matrix. Let us further assume that there exists a $k \times 1$ vector of variables $X(t)$ that are correlated with the first element of $\varepsilon(t)$, which we denote $\varepsilon_1(t)$, and uncorrelated with any other elements of $\varepsilon(t)$.

The question we want to address is how, and under which conditions, $X(t)$ can be used to estimate the impulse response implied by an innovation to the structural disturbance $\varepsilon_1(t)$. As we show below, a condition that allows such identification is that there be only one non-zero element in the first column of Π.[1] This condition can be interpreted as there being an indicator of $\varepsilon_1(t)$ within $y(t)$, that is, the effect of $\varepsilon_1(t)$ on all variables other than the indicator can be viewed as being transmitted through the indicator. Under this *unique indicator* condition, it is easy to transform (2.1) as to highlight how $X(t)$ can be used to estimate the effect of $\varepsilon_1(t)$ on $y(t)$.

[1] It is important to note that this condition is not vacuous since it is not always possible to transform (2.1) such as to make this condition hold.

Without loss of generality, assume that $y_1(t)$ represents the unique indicator of $\varepsilon(t)$, that is, let $y_1(t)$ be chosen such that the only non-zero element in the first column of Π is the first element. Also let A be partitioned as follows:

$$A = \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}, \tag{2.2}$$

where A_{11} is 1×1, A_{12} is $1 \times (n-1)$, A_{21} is $(n-1) \times 1$ and A_{22} is $(n-1) \times (n-1)$. Now define a transformation matrix T as follows:

$$T = \begin{pmatrix} D & DA_{12}(I_{n-1} - A_{22})^{-1} \\ 0_{n-1} & (I_{n-1} - A_{22})^{-1} \end{pmatrix}. \tag{2.3}$$

In Eq. (2.3), I_{n-1} is an identity matrix of order $n-1$, 0_{n-1} is an $(n-1) \times 1$ vector of zeroes and D is a scalar defined by $D = [(1 - A_{11}) - A_{12}(I_{n-1} - A_{22})^{-1}A_{21}]^{-1}$. By pre-multiplying both sides of (2.1) by the matrix T and letting the vector $a^* = \{a_2^*, a_3^*, \ldots, a_n^*\}'$ be defined by $a^* = (I_{n-1} - A_{22})^{-1}A_{21}$, the structural relationship between $y(t)$ and $\varepsilon(t)$ can be rewritten as

$$y(t) = \begin{bmatrix} 0 \\ a_2^* \\ a_3^* \\ \vdots \\ a_n^* \end{bmatrix} y_1(t) + B^*(L)y(t-1) + \Pi^*\varepsilon(t), \tag{2.4}$$

where $B^*(L) = TB(L)$ and $\Pi^* = T\Pi$.

There are two important properties in this transformation. First, it allows one to express the contemporaneous relationship between the elements of $y(t)$ as depending only on $y_1(t)$. Second, and most importantly, the transformation induced by T assures that the first column of Π^* inherits the structure of the first column of Π, that is, if only the first element within the first column of Π is nonzero (as is being assumed), then only the first element within the first column of Π^* is nonzero. This property of Π^* has several implications for the estimation of Eq. (2.4). First, unbiased estimates of a_2^* through a_n^* and $B^*(L)$ can be obtained by using $X(t)$ to instrument $y_1(t)$ in Eq. (2.4). Second, the structural impulse response implied by a change in $\varepsilon_1(t)$ can be calculated by dynamic forecasting where the initial impulse to $y(t)$ corresponds to $(1, a_2^*, a_3^*, \ldots, a_n^*)$ and the dynamics is governed by the estimates of $B^*(L)$. Third, the validity of this identification strategy can be examined by the use of over-identification tests if $k > 1$.

It is important to note that, within this framework, the appropriate interpretation of the over-identification test is as a joint test of hypotheses that $X(t)$ is uncorrelated with shocks other than $\varepsilon_1(t)$ and that $y_1(t)$ is actually the unique indicator of $\varepsilon_1(t)$. Consequently, in the case of the monetary business cycle, it is clear that having access to an instrument that is correlated with only monetary shocks is not generally sufficient to identify the desired impulse response: there is an additional requirement that the monetary policy variable used as an indicator must capture all the direct effect of the monetary innovation. Moreover, the above discussion should make clear that the appropriateness of an instrumental variable strategy depends on a particular system of variables that is under consideration. It may be the case where a measure of monetary policy satisfies the unique indicator condition in one system of variables, but not in another.

Before preceding further, it seems relevant to compare the above procedure with the instrumental variable procedure adopted by Romer and Romer (1990).[2] These authors examine the effect of monetary shocks in a bi-variate version of Eq. (2.4) where they estimate only the second equation and instrument both current and lagged values of the monetary policy variable, which they take to be the growth of M1.[3] They then calculate an impulse response by simulating (using one equation) the effect of a permanent increase in M1. There are three potential drawbacks of such an approach. First, the resulting impulse response is very hard to interpret: it corresponds to the effect of a sequence of unanticipated shocks that ex post happens to trace out the effects of a one time increase in the money supply. This is not the structural impulse response as it is usually defined. Second, the procedure may be inefficient since there is no obvious need to instrument lagged values of the monetary policy variable (conditional on there being enough lags included in the estimation). Finally, and possibly more importantly, Romer and Romer do not indicate or test the conditions under which instrumenting M1 will lead to an unbiased estimate of monetary shocks.

2.1. Evaluating Wold-causal orderings

The above discussion suggests a method for estimating the effects of monetary shocks when instrumental variables are available. This framework can also be

[2] In Romer and Romer (1989), the authors do not actually use their constructed dummy variables to instrument a monetary policy variable. Instead, they estimate a reduced form where the instruments are directly incorporated into prediction equations. Although such a procedure is potentially valid for testing whether monetary shocks have real effects, it is not at all clear whether a meaningful impulse response can be derived from this approach. For example, even if monetary innovations had only a contemporary effect on output, but the lags of the Romer and Romer dummies actually help predict monetary innovations, then the impulse response calculations performed by Romer and Romer (1989) would indicate spuriously that monetary shocks have a persistent effect on output.

[3] Ramey (1993) also uses the same procedure.

used to explore whether alternative identification schemes based on Wold-causal orderings (Choleski decompositions) are appropriate. In fact, under the assumption that the restrictions inherent to a Wold ordering are true, the associated Choleski decomposition provides the most efficient estimation procedure and therefore should be favored over an instrumental variable procedure.

The restrictions implied by a Wold-causal ordering can easily be interpreted within the framework of Eq. (2.4).[4] For example, if $y_1(t)$ is assumed to be at the bottom of a Wold ordering (contemporaneously exogenous), then in terms of Eq. (2.4) it implies that the first row of Π^* has a nonzero element only in the first entry[5] and hence no instrumenting of $y_1(t)$ is necessary. Alternatively, $y_1(t)$ being at the top of a Wold ordering implies that a_2^* through a_n^* are zero and therefore need not be estimated.

In general, a Wold ordering implies that the vector $\tilde{y}(t) = \{y_2(t), \dots, y_n(t)\}$ can be divided into subsets: one set for which the coefficients in a^* are zero (variables that are lower in the ordering) and a second set for which $y_1(t)$ can be considered contemporaneously exogenous (variables that are higher in a Wold ordering). Therefore, the validity of a particular Wold ordering can be tested by examining (i) whether the estimates of a_i^* found by instrumenting $y_1(t)$ are zero for all variables i assumed to be prior to $y_1(t)$ in a Wold ordering, and (ii) whether the coefficients a_i^* obtained by instrumenting $y_1(t)$ are different from those obtained by OLS for all variables i that are assumed to be higher in a Wold ordering. This latter test can be performed by the use of the Hausman–Wu specification test (see Hausman, 1983).

2.2. Extending to the case of cointegration

The above discussion assumes that the statistical model of interest can be written as a vector autoregression (VAR) of stationary variables. In many economic environments, however, this may not be an appropriate assumption. In particular, if the variables of interest have been made stationary by differencing, it is possible that the levels of the variables are cointegrated. In this case, the transformed variables will not have a VAR representation as assumed in Eq. (2.1). Instead the system will have a vector error correction (VEC) representation as given in Eq. (2.5), where the additional vector $z(t)$ in Eq. (2.5) is

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[4] For a discussion of the restriction underlying Choleski decompositions within the setting of Eq. (2.1), see Hausman (1983) or Bernanke (1986).

[5] In order for (2.4) to be estimated consistently by OLS, $\Pi_{12} + A_{12}(I_{n-1} - A_{22})^{-1}\Pi_{22}$ must happen to be $0'_{n-1}$, where $\Pi = \begin{pmatrix} \Pi_{11} & \Pi_{12} \\ 0_{n-1} & \Pi_{22} \end{pmatrix}$.

a $r\times 1$ vector of cointegrating relationships ($r<n$) and C is a $n\times r$ matrix.[6]

$$y(t)=Ay(t)+B(L)y(t-1)+Cz(t-1)+\Pi\varepsilon(t). \tag{2.5}$$

The case with cointegration can nevertheless be handled in a manner identical to that of non-cointegration. In fact, by pre-multiplying Eq. (2.5) by the matrix T defined in Eq. (2.3) we obtain an analogue to Eq. (2.4) with the sole exception of a need to include cointegrating relationships as additional regressors. Hence, the previous strategy for estimating the effects of monetary shocks, as well as evaluating the plausibility of Wold orderings, can be carried out as described above.

3. Estimation results

3.1. Comparing different identification schemes

There are two remaining issues to discuss before the above framework can be used to examine the effects of monetary shocks. First, there is a choice of indicators for monetary policy. Our reading of the literature suggests examining at least three measures: M1, non-borrowed reserves[7] and the Federal Funds rate. Second, there is a choice of potential instrumental variables. For this purpose we use the dummy variable constructed by Romer and Romer (1989), and its lags, as our set of instruments.[8]

All of our estimations are carried out with monthly data. We choose the dimension of $y(t)$ to be 4, with the raw variables being the total index of industrial production, the consumption price index, the three month T-bill rate (auction average) and one of the indicators of monetary policy. The series are taken from CITIBASE. The adopted sample period is from January 1959 to December 1987. This choice reflects the fact that the non-borrowed reserves series start from 1959 and that Romer and Romer examine the historical record only up to 1987. During the sample period, Romer and Romer identify four episodes as monetary contractions: December 1968, April 1974, August 1978 and October 1979. The Romer dummy variable takes on the value of one in these months and takes the value of zero otherwise.

[6] See Watson (1994) for an overview of cointegration literature.

[7] Strongin (1995) suggests using the ratio of non-borrowed to total reserves as a measure of monetary policy. We found that this measure performed almost identically to the growth of non-borrowed reserves and therefore omitted the results.

[8] By including the lags of the Romer dummy variable, we implicitly allow for the date identified as monetary contractions to be indicators of future monetary shocks.

In order to be careful with respect to the order of integration of variables, we considered several transformations of the data. Our base case corresponds to a VAR representation (as in Eq. Eq. (2.4)) where, when used, output, non-borrowed reserves and M1 are in growth rates, while the T-bill rate, the inflation rate and the Federal Funds rate are in differences. We present this specification as our base case since, under this transformation, the data appear stationary (when tested using augmented Dickey–Fuller tests). As indicated previously, however, a VAR representation for this system may not be appropriate if there are cointegrating relationships among the variables. Therefore, we also report results based on a VEC representation under different assumptions regarding cointegrating relationships. It is immediately worth noting that we found our results to be very robust with respect to alternative assumptions regarding integration and cointegration of variables.

We estimate each four-variables system of equations by both OLS and IV. We conduct Hausman's specification test to examine whether the contemporaneous coefficients on the monetary policy variable are sensitive to the use of instruments. Since the validity of the instrumental variable procedure is obviously questionable, we also report Basmann's over-identification test statistics (see Basmann, 1960).

Tables 1, 3 and 5 report, for different lag length of $B^{*}(L)$, the OLS and IV estimates of a_i^{*} ($i = 2$ to 4), the Hausman statistics and the over-identification test statistics corresponding to our base case specification. Although both the Akaike information criterion and the likelihood ratio test generally indicate that 12 lags in $B^{*}(L)$ are preferred to 6 and 18, the absence of perfect conformity in these results leads us to report all three cases. The number of lags of the Romer dummy variable is set at 24. Results are very similar if we use 6, 12 or 18 lags of this dummy variable, with the exception that the estimates of a_i^{*} become slightly less precise as the number of lags decreases.[9]

In Table 1, the indicator of monetary policy that plays the role of $y_1(t)$ is the aggregate of non-borrowed reserves. The first observation to take from this table is that there is very little evidence against the instrumental variable procedure. In fact, there is no evidence against the underlying identifying restrictions at the 1% level, and there is only one case out of nine with rejection at the 10% level. Therefore, we consider acceptable the joint hypothesis that the Romer and Romer dummies are correlated only with monetary shocks and that the aggregate of non-borrowed reserves is a good indicator of monetary policy.

The second inference that can be made from Table 1 is that any Choleski decomposition imposed on this system of four variables is likely to provide biased estimates of the effects of monetary shocks. The problem with using

[9] Romer and Romer (1989) use 36 lags of this variable in their analysis.

Table 1
Non-borrowed reserves as measure of monetary policy: system of $\Delta\ln(NBR)$, $\Delta(TB\ rate)$, $\Delta\ln$ (*Industrial Production*) and Δ(*Inflation*)

Dependent variable	Number of lags[a]	Non-instru-mented[b] (OLS)	Instru-mented[b] (IV)	Hausman–Wu test[c], *t*-value	Over-identification test for R&R dummies[d]
Change in treasury bill rate	6	− 0.00810	− 0.02206	3.52091	1.30258
		(0.00113)	(0.00412)		(0.15611)
	12	− 0.00813	− 0.02273	3.92860	1.31342
		(0.00114)	(0.00389)		(0.15050)
	18	− 0.00733	− 0.01743	3.07054	1.65833
		(0.00115)	(0.00348)		(0.02931)
Growth of industrial production	6	− 0.03847	− 0.12812	1.28257	1.24652
		(0.02419)	(0.07396)		(0.19727)
	12	− 0.02578	− 0.07143	0.78217	1.21959
		(0.02288)	(0.06269)		(0.22119)
	18	− 0.05486	− 0.15711	1.67892	1.01109
		(0.02383)	(0.06540)		(0.45345)
Change in inflation	6	− 0.01204	0.01071	− 1.10536	0.87783
		(0.00717)	(0.02180)		(0.63675)
	12	− 0.01019	− 0.00204	− 0.44747	0.99610
		(0.00718)	(0.01958)		(0.47271)
	18	− 0.00990	− 0.00150	− 0.45438	0.93989
		(0.00751)	(0.01996)		(0.55002)

[a]The number of lags for R&R dummies is set at 24 in all cases.
[b]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parentheses are standard errors.
[c]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The *t*-value is reported.
[d]The value of F tests is reported. The numbers in parentheses show the level of significance.

a Choleski decomposition to identify the effects of ε_1 on interest rates is that it is incompatible with a two-way interaction between interest rates and non-borrowed reserves. The Hausman specification test, however, suggests that non-borrowed reserves are correlated with innovations in interest rates, and the IV estimates of the coefficients on non-borrowed reserves in the interest rate equations strongly suggest that innovations in non-borrowed reserves affect the T-bill rate, that is, there is a clear indication of a two way interaction. Therefore, contrary to the suggestion of Christiano and Eichenbaum (1992) and Strongin (1995), using non-borrowed reserves as an indicator of monetary policy does not seem to overcome the endogeneity problem usually associated with more inclusive measures of money.

Table 2
Non-borrowed reserves as measure of monetary policy: alternative specifications (number of lags = 12)

Dependent variable	Non-instrumented[b] (OLS)	Instrumented[a] (IV)	Hausman–Wu Test[b], *t*-value	Over-identification test for R&R dummies[c]
Panel A: System of $\Delta \ln(NBR)$, $\Delta(TB\ rate)$, $\Delta \ln(Industrial\ Production)$ and $\Delta(Inflation)$ with cointegration[d]				
$\Delta(TB)$	− 0.00767	− 0.02142	3.71248	1.35184
	(0.00113)	(0.00387)		(0.12736)
$\Delta \ln(IP)$	− 0.02429	− 0.08390	0.98182	1.25302
	(0.02310)	(0.06497)		(0.19367)
$\Delta(Inflation)$	− 0.01032	− 0.00282	− 0.39776	0.99425
	(0.00725)	(0.02020)		(0.47522)
Panel B: System of $\Delta \ln(NBR)$, $(TB\ rate)$, $\Delta \ln(Industrial\ Production)$ and $(Inflation)$				
(TB)	− 0.00755	− 0.02138	3.60622	1.36240
	(0.00113)	(0.00400)		(0.12141)
$\Delta \ln(IP)$	− 0.02761	− 0.08566	0.93408	1.08308
	(0.02292)	(0.06625)		(0.36196)
$(Inflation)$	− 0.01106	− 0.00322	− 0.40797	0.95330
	(0.00717)	(0.02053)		(0.53121)
Panel C: System of $\Delta \ln(NBR)$, $(TB\ rate)$, $\Delta \ln(Industrial\ Production)$ and $\Delta(Inflation)$ with cointegration[e]				
(TB)	− 0.00777	− 0.02211	3.66392	1.26921
	(0.00114)	(0.00408)		(0.18123)
$\Delta \ln(IP)$	− 0.03253	− 0.08398	0.82376	1.13425
	(0.02307)	(0.06659)		(0.30399)
$(Inflation)$	− 0.01111	− 0.00477	− 0.32585	0.95485
	(0.00725)	(0.02076)		(0.52907)

[a]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parentheses are the standard errors.
[b]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The *t*-value is reported.
[c]The value of F tests is reported. The numbers in parentheses are the level of significance. As in Table 1, the number of lags for R&R dummies is set at 24 in all cases.
[d]The one-period lagged value of $[TB\ rate - Inflation]$ is included in this specification.
[e]The one-period lagged value of the co-integration vector estimated by the dynamic OLS is included in this specification ($\ln(NBR/P) + 0.034(0.027) \times \ln(Industrial\ Production) + 77.0(3.61) \times TB\ rate$).

Table 3 reports results for the case where M1 is used as a measure of monetary policy. One way to interpret the results of this table is as a test of the power of our approach. In general, it is considered that M1 is not a very good indicator of monetary policy. Therefore, if our method of inference has any

Table 3
M1 as measure of monetary policy: system of $\Delta\ln(M1)$, $\Delta(TB\ rate)$, $\Delta\ln(Industrial\ Production)$ and $\Delta(Inflation)$

Dependent variable	Number of lags[a]	Non-instrumented[b] (OLS)	Instrumented[b] (IV)	Hausman–Wu test[c], t-value	Over-identification test for R&R dummies[c]
Change in treasury bill rate	6	– 0.01050 (0.00591)	– 0.01228 (0.01808)	0.10476	3.09812 (0.00000)
	12	– 0.00524 (0.00615)	– 0.01745 (0.01868)	0.69170	3.34912 (0.00000)
	18	– 0.00342 (0.00608)	– 0.01456 (0.01771)	0.66967	2.94113 (0.00001)
Growth of industrial production	6	– 0.12947 (0.12569)	– 0.51487 (0.38993)	1.04412	1.32274 (0.14306)
	12	– 0.00913 (0.12379)	– 0.29883 (0.37725)	0.81296	1.35287 (0.12672)
	18	0.04573 (0.12964)	– 0.27557 (0.37952)	0.90080	1.26169 (0.18878)
Change in inflation	6	– 0.02093 (0.03726)	0.06810 (0.11494)	– 0.81877	0.84632 (0.68024)
	12	– 0.02585 (0.03797)	0.04668 (0.11534)	– 0.66602	0.96272 (0.51819)
	18	– 0.04602 (0.03925)	– 0.01305 (0.11368)	– 0.30902	0.83274 (0.69764)

[a]The number of lags for R&R dummies is set at 24 in all cases.
[b]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parentheses are the standard errors.
[c]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The t-value is reported.
[d]The value of F tests is reported. The numbers in parentheses show the level of significance.

power, we should reject the use of M1 as a means of identifying monetary shocks. The result for the over-identification test supports this conjecture: the rejection of the over-identifying restrictions suggests that innovations in monetary shocks affect interest rates through channels not captured by M1.

Our most controversial results are probably in Table 5, where it is the Federal Funds rate that is used as an indicator of monetary policy. In this case, as in the case with M1, we again find strong evidence against the over-identification restriction. This places in doubt Bernanke and Blinder's (1992) conclusion that the Federal Funds rate is a good indicator of monetary policy. It is, however, important to be careful in interpreting these results. The simplest interpretation is that a monetary contraction (or expansion) does not only work through the level of the Federal Funds rate, but it also works through the contemporaneous spread between the Federal Funds rate and the T-bill rate; for this reason, the Federal Funds rate is not a good measure of monetary policy since it is not capturing the full effect of a monetary shock as is required to satisfy the unique indicator condition discussed in Section 2. Therefore, even if the Federal Funds rate is contemporaneously exogenous with respect to innovations in money demand, as suggested by Bernanke and Blinder (1992), it may still be a measure of monetary policy that is inferior to non-borrowed reserves.

In summary, the evidence presented in Tables 1, 3 and 5 rejects the use of either M1 or the Federal Funds rate as an indicator of monetary policy, but does not reject the use of non-borrowed reserves. Moreover, the evidence runs counter to the view that non-borrowed reserves can be used in conjunction with a Choleski decomposition to identify the effects of monetary shocks. Therefore, it seems that a combined procedure of instrumenting non-borrowed reserves with the Romer dummy variable may provide the most defensible means of identifying the effects of monetary shocks. However, before rushing to conclusions, it seems warranted to further examine the robustness of these results.

Tables 2, 4 and 6 examine the robustness of the above inferences with respect to different assumptions regarding cointegrating relationships. Since it is difficult to identify the exact nature and number of cointegrating relationships in a system (due to the low power of tests in small samples), we report a set of results based on different theoretically plausible views. Panel A of Table 2 reports the results analogue to those presented in Table 1 (based on 12 lags) with the exception that the system of equations is now estimated under the assumption that the nominal interest rate and the inflation rate are cointegrated with cointegrating vector $(1, -1)$. In other words, the system of equations is estimated with the inclusion of a lagged real interest rate as an additional regressor. In Panel B, the nominal interest rate and the inflation rate are assumed to be stationary and no cointegration relationship is imposed. Finally Panel C assumes, as in Panel B, that the nominal interest rate and the inflation rate are stationary but imposes the additional assumption of a money demand cointegrating relationship.[10] In this case the cointegrating vector between output, real money, and the nominal interest rate is estimated by the dynamic ordinary least-squares method suggested by Stock and Watson (1993). The main aspect to notice from all three of these Panels is the extreme similarity with the results presented in Table 1. In particular, the evidence supports the view that the aggregate of non-borrowed reserves is a good indicator of monetary policy, but nevertheless cannot be placed within a Wold causal ordering. This observation is not too surprising since taking into account cointegrating relationships is

[10] The plausibility of such a cointegrating relationship is emphasized in King et al. (1991).

Table 4
M1 as measure of monetary policy: alternative specifications (number of lags = 12)

Dependent variable	Non-instrumented[a] (OLS)	Instrumented[a] (IV)	Hausman–Wu test[b], t-value	Over-identification test for R&R dummies[c]
Panel A: System of $\Delta\ln(M1)$, Δ (*TB rate*) $\Delta\ln$(*Industrial Production*), and Δ(*Inflation*) with cointegration[d]				
Δ (*TB*)	− 0.00613	− 0.02621	1.14184	2.90041
	(0.00600)	(0.01858)		(0.00001)
$\Delta\ln(IP)$	− 0.01310	− 0.34617	0.92438	1.45119
	(0.12394)	(0.38104)		(0.08067)
Δ(*Inflation*)	− 0.02606	0.04517	− 0.64837	0.97005
	(0.03806)	(0.11626)		(0.50812)
Panel B: System of $\Delta\ln(M1)$, (*TB rate*), $\Delta\ln$(*Industrial Production*) and (*Inflation*)				
(*TB*)	− 0.00621	− 0.03131	1.39168	2.82109
	(0.00604)	(0.01902)		(0.00002)
$\Delta\ln(IP)$	0.03683	− 0.22204	0.72740	1.33388
	(0.12214)	(0.37627)		(0.13770)
(*Inflation*)	− 0.01882	0.07623	− 0.87103	0.81935
	(0.03731)	(0.11532)		(0.71600)
Panel C: System of $\Delta\ln(M1)$, (*TB rate*), $\Delta\ln$(*Industrial Production*) and (*Inflation*) with Co-integration[e]				
(*TB*)	− 0.00599	− 0.02865	1.26396	2.86894
	(0.00610)	(0.01894)		(0.00001)
$\Delta\ln(IP)$	0.02702	− 0.19131	0.61618	1.41540
	(0.12217)	(0.37479)		(0.09540)
(*Inflation*)	− 0.01799	0.06874	− 0.79564	0.84092
	(0.03743)	(0.11525)		(0.68717)

[a]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parenthses are the standard errors.
[b]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The t-value is reported.
[c]The value of F tests is reported. The numbers in parentheses are the level of significance. As in Table 3, the number of lags for R&R dummies is set at 24 in all cases.
[d]The one-period lagged value of [*TB rate* - *Inflation*] is included in this specification.
[e]The one-period lagged value of the cointegration vector estimated by the dynamic OLS is included in this specification ($\ln(M1) - \underset{(0.050)}{0.373} \times \ln(\textit{Industrial Production}) + \underset{(4.87)}{46.0} \times \textit{TB rate}$).

Table 5
Federal Funds rate as measure of monetary policy: system of Δ(*FF rate*), Δ(*TB rate*), Δln(*Industrial Production*), and Δ(*Inflation*)

Dependent variable	Number of lags[a]	Non-instrumented[b] (OLS)	Instrumented[b] (IV)	Hausman–Wu test[c], *t*-value	Over-identification test for R&R dummies[d]
Change in treasury bill rate	6	0.60610	0.69647	− 1.59613	2.50550
		(0.03387)	(0.06597)		(0.00015)
	12	0.58639	0.70184	− 2.03991	3.08693
		(0.03566)	(0.06689)		(0.00000)
	18	0.57145	0.66790	− 1.75942	2.10461
		(0.03575)	(0.06545)		(0.00233)
Growth of industrial production	6	3.89866	2.66335	0.77085	1.50138
		(0.97005)	(1.87326)		(0.06201)
	12	3.39046	3.34402	0.03206	1.25353
		(0.93633)	(1.72499)		(0.19321)
	18	4.00757	4.80710	− 0.53960	1.10811
		(0.98429)	(1.77884)		(0.33423)
Change in inflation	6	0.35825	0.86322	− 1.04822	0.70104
		(0.29073)	(0.56268)		(0.85554)
	12	0.37117	0.85844	− 1.05875	0.91845
		(0.29545)	(0.54690)		(0.57981)
	18	0.39473	0.96722	− 1.16765	0.78952
		(0.32344)	(0.58736)		(0.75372)

[a]The number of lags for R&R dummies is set at 24 in all cases.
[b]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parenthses are the standard errors.
[c]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The *t*-value is reported.
[d]The value of *F* tests is reported. The numbers in parentheses show the level of significance.

mainly helpful in capturing long-run behavior, while the focus of our tests rests on short-run behavior.

Tables 4 and 6 are similar in spirit to Table 2. In fact, Table 4 reports the results exactly analogous to those in Table 2 with the exception that the indicator of monetary policy is now M1 instead of non-borrowed reserves. The results from Table 4 essentially replicate those reported in Table 3, thereby providing further support for the view that M1 is not an appropriate indicator of monetary policy. The cases considered in Table 6 are slightly different from those of Tables 2 and 4 since the most plausible cointegrating relationships are different. For example, in Panel A of Table 6 the maintained assumptions are (i) that the federal funds rate and the T-bill rate are cointegrated with vector (1, − 1) and

Table 6
Federal Funds rate as measure of monetary policy: alternative specifications (number of lags = 12)

Dependent variable	Non-instrumented[a] (OLS)	Instrumented[a] (IV)	Hausman–Wu Test[b], t-value	Over-identification test for R & R dummies[c]
Panel A: System of $\Delta(FF)$, $\Delta(TB\ rate)$, $\Delta\ln(Industrial\ Production)$, and $\Delta(Inflation)$ with Co-integration[d]				
$\Delta\ (TB)$	0.57180	0.67027	− 1.71601	3.00116
	(0.03552)	(0.06748)		(0.00001)
$\Delta\ln(IP)$	3.43713	3.81791	− 0.25815	1.12316
	(0.92981)	(1.74362)		(0.31613)
$\Delta(Inflation)$	0.38538	0.94488	− 1.16900	0.88933
	(0.29925)	(0.56446)		(0.62051)
Panel B: System of $(FF\ rate)$, $(TB\ rate)$, $\Delta\ln(Industrial\ Production)$, and $(Inflation)$				
(TB)	0.58404	0.68449	− 1.72719	3.24907
	(0.03562)	(0.06819)		(0.00000)
$\Delta\ln(IP)$	3.62456	3.80396	− 0.11940	1.07929
	(0.93783)	(1.77122)		(0.36648)
$(inflation)$	0.43758	1.10332	− 1.38111	0.80030
	(0.29728)	(0.56633)		(0.74078)

[a]The estimated contemporaneous parameter on the monetary policy measure is reported. The numbers in parentheses are the standard errors.
[b]The Hausman–Wu statistics test the equality between the OLS and IV estimates. The t-value is reported.
[c]The value of F tests is reported. The numbers in parentheses are the level of significance. As in Table 5, the number of lags for R&R dummies is set at 24 in all cases.
[d]The one-period lagged values of both [$TB\ rate - Inflation$] and [$TB\ rate - FF\ rate$] are included in this specification.

(ii) the T-bill rate and the inflation rate are also cointegrated with vector (1, − 1). In contrast, Panel B of Table 6 does not impose any cointegrating relationships but does instead assume that the federal funds rate, the T-bill rate and the inflation rate are all stationary in levels. Again, the results from Table 6 are extremely similar to those reported in Table 5. Consequently, we conclude that the inferences drawn from Tables 1, 3 and 5 are robust with respect to different assumptions regarding the order of integration and cointegration of variables.

3.2. Examining the power of the approach

Our empirical investigation suggests that non-borrowed reserves satisfy the unique indicator condition discussed in Section 2, and can therefore potentially be used to identify the quantitative effect of monetary shocks. We inferred this result from the observation that none of the test statistics associated with the over-identifying restrictions in either Table 1 or Table 2 are rejected. It is, however, reasonable to question the appropriateness of this inference since it is known that the Basmann over-identification test may have low power (see Newey, 1985). In order to evaluate this possibility, we follow two routes. First we examine whether our testing procedure would lead us to infer that other measures of money may also satisfy the unique indicator condition. In particular, if the power of our testing procedure is very low, we would expect to accept this null often. Then we report results from a set of Monte Carlo experiments aimed at illustrating the property of the Basmann over-identification test in our setup.

Table 7 reports the over-identification test statistics associated with the use of different measures of money.[11] The reported statistics correspond to the test of whether the error in the T-bill equation is correlated with the instruments (Basmann's test). We only report the results for the T-bill equation since, as in Tables 1–6, this test is never rejected based on either the output or inflation equations. The pattern of results in Table 7 is clear. There is strong evidence against the hypothesis that any of the measures of money, except for non-borrowed reserves, satisfies the unique indicator condition. Although this pattern of results does not prove that our testing procedure has power, the fact that we accept the unique indicator hypothesis only in a case which is plausible on a priori grounds gives credibility to the result.

In order to more directly examine the property of our over-identification test, we conducted two sets of Monte Carlo experiments. In both cases, the data generating process is estimated from our data using the specifications adopted in Table 1 and imposing the restrictions implied by either the null or the alternative hypothesis. Our first experiement is aimed at examining size distortions under the hypothesis that the monetary indicator satisfies the unique indicator condition and that the instruments are only correlated with monetary shocks. The second experiment examines the power of the test under an alternative hypothesis where the monetary indicator does not satisfy the unique indicator condition. In particular, under the alternative it is assumed that the first column of the Π matrix has all nonzero entries except in its first entry. Table 8 reports the result for these two experiments.

The first aspect to notice from Table 8 is that the size distortions are not very substantial with the possible exception of the T-bill equation. In the case of the T-bill equation, the Monte Carlo results suggest that the test may be being rejected too often. However, if we reexamine our rejections in Tables 1–7 in the

[11] The data on monetary aggregates is again from Citibase and each VAR is estimated over the period from January 1959 to December 1987.

Table 7
Over-identification tests for alternative monetary aggregates[a]

Case of $\Delta(TB\text{-}rate)$ equation[b]			
Number of lags	6	12	18
Non-borrowed reserves	1.30258	1.31342	1.65833
	(0.156)	(0.151)	(0.029)
Total reserves	2.36049	3.21638	3.03613
	(0.000)	(0.000)	(0.000)
Monetary base	2.47186	3.35231	3.04791
	(0.000)	(0.000)	(0.000)
M1	3.09812	3.34912	2.94113
	(0.000)	(0.000)	(0.000)
M2	2.17256	2.30608	2.88404
	(0.001)	(0.001)	(0.000)
M3	3.66727	4.09870	3.41511
	(0.000)	(0.000)	(0.000)
M3 plus other liquid assets	3.44518	3.82544	2.93960
	(0.000)	(0.000)	(0.000)

[a]The value of F tests is reported. The numbers in parentheses are the level of significance.
[b]The system consists of growth of monetary aggregates, changes in T-bill rates, growth of industrial production, and changes in inflation rates. Only the results for the over-identification test associated with the T-bill equation are reported.

light of this size distortion, we are not led to change any of our inferences since all rejections were associated with extremely low p-values. Hence, potential size distortions do not appear to place in doubt our results. Secondly, when we look at the property of this test under a hypothesized violation of the unique indicator condition, it still appears to have reasonable power in the case of the T-bill equation. For example, even when correcting for size distortions, we find that the power of this test is not reduced substantially (compare the sixth column with the fifth in Table 8), and that the test is still rejected more than 50% of the experiments under the alternative hypothesis. In summary, we take the results of Table 8 as indicating that this over-identification test has decent properties in our setup.

3.3. Application: Estimating the liquidity effect

Recently there has been a considerable interest in identifying the effect on interest rates of monetary shocks. In particular, Christiano and Eichenbaum (1992) and Strongin (1995) have documented the extent to which interest rates fall after an innovation in non-borrowed reserves and have presented their results as supporting the view of strong liquidity effects. However, given the

Table 8
Monte Carlo study of properties of over-identification tests[a]

Dependent variable	Number of lags	Size of test[b] (5%)[c]	Size of test[b] (10%)	Power of test[d] (10%)	Size corrected power (10%)
Change in	6	8.9	15.4	73.4	67.5
treasury bill	12	10.1	17.4	60.0	51.3
rate	18	7.5	12.9	73.8	64.2
Growth of	6	3.2	7.1	90.4	91.7
industrial	12	4.0	7.9	90.9	92.8
production	18	4.0	8.6	80.0	83.6
Change in	6	3.5	6.7	81.8	87.6
inflation	12	3.4	6.9	81.1	85.6
	18	3.8	8.2	66.8	73.8

[a]The data generating process used to perform each Monte Carlo experiment was inferred from the system of $\Delta\ln(NBR)$, $\Delta(TB\ rate)$, $\Delta\ln(Industrial\ Production)$, and $\Delta(Inflation)$ under either the null hypothesis or the alternative hypothesis. A different data generating process was estimated for each lag length. 5000 simulations were used to calculate rejection rates.
[b]Entries in these columns represent the percentage of times that the statistics are rejected when the critical value is chosen such that a rejection rate of F-statistics would be 5% or 10% under the null.
[c]The size of the test is calculated under the null hypothesis that the monetary indicator satisfies the unique indicator condition and that the instrument is only correlated with the monetary shock.
[d]The power of the test is calculated under the alternative hypothesis that the monetary indicator does not satisfy the unique indicator condition.

evidence of the previous section, the procedure followed by these authors is likely to have provided biased estimates of the effect of a monetary innovation. In order to quantify this potential bias and provide an example of our estimation strategy, Fig. 1 compares the impulse response obtained by our instrumental variable approach with that obtained by the use of a Choleski decomposition (which is the identification method favored by both Christiano and Eichenbaum and Strongin).

In Fig. 1, the dotted line represents the impulse response associated with identifying monetary injections with innovations in non-borrowed reserves, where the aggregate of non-borrowed reserves is placed after the three-month T-bill rate and before output growth and inflation in a Wold-causal ordering. The variables and the sample are the same as those used in the previous section. The number of lags used in this estimation is 12. The full heavy line represents the impulse response associated with estimating Eq. (2.4) when 24 lags of the Romer dummy variable are used to instrument the contemporaneous value of non-borrowed reserves. The important element to note from this figure is that the failure to take into account the endogeneity of non-borrowed reserves with

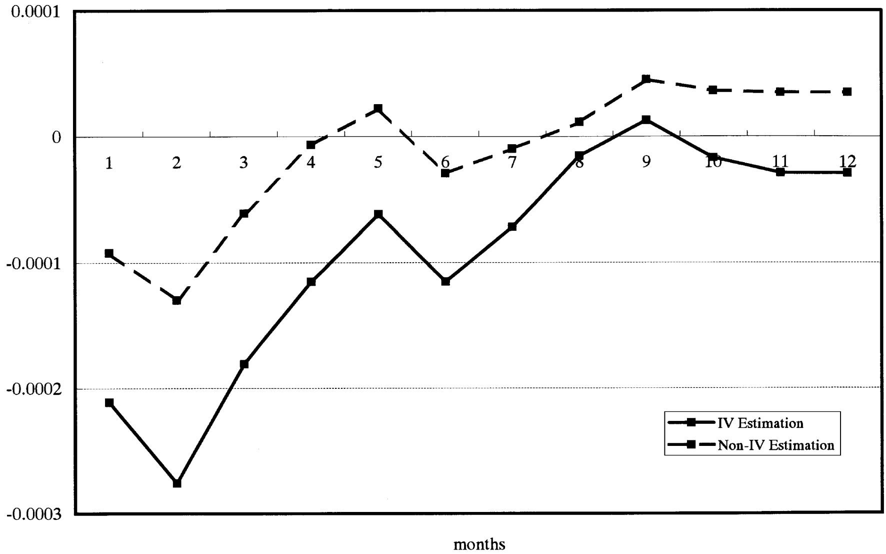

Fig. 1. Effects on treasury bill rates.

respect to the interest rate is likely to under-estimate the liquidity effect. In effect, the point estimates for the first several months following monetary innovations are almost twice as large when the IV procedure is used as compared to a Choleski decomposition (the Hausman statistics reported in Tables 1 and 2 indicate that this difference is statistically significant). The direction of this bias is not surprising since it is the same direction as the one found when replacing in a Choleski decomposition a more inclusive measure of money, like M1, with a more restrictive measure, like non-borrowed reserves.

4. Conclusion

Given the scarcity of legitimate instrumental variables in empirical macroeconomics, the historical episodes identified as monetary contractions by Romer and Romer (1989) provide researchers with one precious candidate that deserves to be exploited fully. Several previous studies have attempted to use the Romer dummies as instruments, however their procedures have not been precisely defined or tested. In contrast, this paper clarifies the condition under which such episodes can be used to identify the effects of monetary shocks and further exploits the framework to test the validity of commonly used Wold orderings. The results show that Wold orderings are generally inappropriate to estimate the effects of monetary shocks on market interest rates and that a procedure consisting of instrumenting non-borrowed reserves with the Romer dummies may be appropriate. Obviously, this framework can be extended to test restrictions imposed by other types of decompositions. Such an extension is left for future research.

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