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Monetary policy and the term structure of nominal interest rates: evidence and theory*

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Abstract

This paper explores how exogenous impulses to monetary policy affect the yield curve for nominally risk-free bonds. Three distinct identification strategies imply similar patterns: a contractionary policy shock induces a pronounced positive but short-lived response of short-term interest rates. The response declines monotonically with maturity; long-term rates are virtually unaffected. These responses are unambiguously liquidity effects rather than expected inflation effects. Monetary-policy shocks account for a relatively small fraction of the long-run variance of interest rates. We find that a limited participation model of monetary nonneutrality is broadly consistent with these empirical patterns.

1 Introduction

Monetary policy is the natural starting point for an inquiry into the economic determinants of the nominal term structure. Bond-traders and other

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nonacademic observers often cite monetary policy as a major factor in term structure movements.[1] Academic observers have also argued that the term structure is intimately linked to monetary policy and its goals. For example, Bernanke and Blinder (1992), Estrella and Hardouvelis (1991), and Mishkin (1990) explore using the spread between long-term and short-term yields as an indicator of monetary policy, future economic activity, and future inflation. However, empirical models of the term structure typically used in the finance literature do not explicitly incorporate monetary policy. Rather, they characterize the nominal term structure as driven by unobserved latent factors. For example, Litterman and Scheinkman (1991) and Dai and Singleton (1997) estimate three-factor models, with the factors associated with the level, slope, and curvature of the yield curve. An open question is whether one or more of these factors corresponds, in part, to monetary policy shocks.

In this paper, we ask how exogenous impulses to monetary policy affect yields on zero-coupon bonds of various maturities. We investigate the impact of these shocks on the shape of the yield curve, as well as on term premiums, and ex ante real rates. Having documented these empirical patterns, we ask if a dynamic stochastic equilibrium model of aggregate economic activity can replicate the empirical patterns we find in the data. Standard equilibrium macroeconomic models have had little success in modeling the term structure.[2] We ask whether the performance of this class of models can be improved by incorporating explicit nominal rigidities. Such an inquiry is a critical step in matching financial factors with economic determinants.

The fundamental empirical problem in assessing the effects of monetary policy shocks is the identification problem: how to distinguish exogenous monetary-policy shocks from the endogenous response of the monetary policy instrument to other, nonmonetary, exogenous impulses. There is disagreement in the profession on the best way to resolve this identification problem. Rather than taking a stand in this controversy, we use three different identification strategies that have been proposed in the literature. All are variants of the identified vector autoregression (VAR) approach proposed by Sims (1986), Bernanke (1986), and Blanchard and Watson (1986). In particular, we use the recursive identification strategy of Christiano, Eichenbaum, and Evans (1996a, b), the nonrecursive identification strategy advocated by Sims and Zha (1995a), and the approach of Gali (1992) that utilizes long-run restrictions as part of the identification strategy.

While these three identification strategies have differing implications for

[1]For example, the *Wall Street Journal* of December 13, 1995 describes February 1994 as the month "when the Fed began raising short-term interest rates and set off the year's bond-market slaughter."

[2]See, for example, den Haan (1995), Backus, Gregory, and Zin (1989), and Bekaert, Hodrick, and Marshall (1997b).

the effect of monetary policy shocks on real economic variables, it is interesting that their implications for the effect of monetary policy shocks on the term structure are broadly similar. All three strategies imply that a contractionary policy shock induces a pronounced positive but transitory response in short-term interest rates, with a smaller effect on medium-term rates and almost no effect on long-term rates. This finding stands in contrast to the popular opinion, often expressed in the financial press, that changes in monetary policy systematically affect long-term bond prices.[3] Our empirical results imply that the main effect of monetary policy shocks is to shift the slope of the yield curve. Because of their transitory impact, monetary-policy shocks account for a relatively small fraction (less than 15%) of the long-run variance of interest rates. This shock roughly corresponds to the slope factor in the models of Litterman and Scheinkman (1991), Knez, Litterman, and Scheinkman (1994), and Dai and Singleton (1997). The response of the yield curve to a monetary-policy shock is unambiguously a liquidity effect rather than an expected inflation effect. That is, the response of the nominal interest rate to a policy shock is in the opposite direction to the response of expected future inflation. We find some evidence that a contrctionary policy shock increases term premiums, at least for shorter maturities.

Having documented these empirical patterns, we ask whether they are consistent with a dynamic stochastic equilibrium model that incorporates nominal rigidities. We focus on the limited participation model suggested by Lucas (1990), Fuerst (1992), and Christiano and Eichenbaum (1995). We calibrate the money-growth process to the results from our estimated VARs. The theoretical model captures the broad features found in the data. In particular, a contractionary monetary shock causes a short-lived rise in the short-term yields, with the response decreasing in the maturity of the bond. These responses are liquidity effects, with the real yield rising substantially more than the nominal yields. In addition, the monetary contraction induces a rise in term premiums, which also decreases with maturity.

A number of recent studies are related to our empirical analysis. Each of our identification strategies has the property that the monetary authority does not respond to developments in the bond market contemporaneously. The studies by Leeper, Sims, and Zha (1996) and Bernanke, Gertler, and Watson (1997) also maintain this assumption in their analysis with both short- and long-term interest rates. While the focus of the latter article is on the way the monetary authority's response function amplifies nonmonetary impulses, they report a number of results that are qualitatively similar to our findings. Gordon and Leeper (1994) and McCallum (1994a), however,

[3] Of course, our empirical experiment has a precise definition. Statements about monetary policy in the financial press undoubtedly confound monetary-policy shocks and normal responses of policy to nonmonetary shocks.

take the view that the monetary authority responds contemporaneously to information conveyed in long-term interest rates.

The plan of the remainder of the paper is as follows: In Section 2 we describe the three strategies we use to identify monetary-policy shocks. In Section 3, we present the implications of each of these strategies for the effect of monetary-policy shocks on the yield curve. Section 4 sets out the equilibrium model with limited participation constraints, describes our calibration of the model, and compares the implications of the theoretical model to our empirical results. Section 5 concludes.

2 Identifying monetary-policy shocks

Since Sims (1980), numerous proposals have been made for identifying fundamental economic impulses using VAR methods.[4] In an attempt to characterize the facts about monetary policy and the term structure robustly, we use three alternative strategies for identifying monetary-policy shocks. Each of the three strategies requires estimation of an identified VAR. To conserve space, the discussion in this section focuses on identification issues. For a discussion of econometric issues in estimating these models, the reader is referred to Christiano, Eichenbaum, and Evans (1997b) and Sims and Zha (1995b).

The empirical approaches we use are: (1) a recursive strategy studied by Christiano, Eichenbaum, and Evans (1996a, b); (2) a nonrecursive strategy studied by Sims and Zha (1995a); and (3) a strategy which employs a combination of long-run and contemporaneous restrictions studied by Gali (1992). Much of the literature focuses on quarterly time series analysis, while the frequency of our data analysis is monthly. Consequently, the robust macroaggregate responses to the three measures of monthly monetary-policy shocks are of some independent interest.

2.1 *Monetary-policy rules*

In all of the identification strategies we use, it is assumed that the monetary policy instrument is the Federal funds rate, denoted FF_t. We assume that FF_t is determined by a relationship of the form

$$FF_t = f(\Omega_t) + \sigma\epsilon_t \tag{1}$$

In equation (1), Ω_t is the information set available to the monetary authority at date t, f is a linear function that describes the monetary authority's

[4]The articles mentioned in the introduction are a small subset of the empirical literature that uses VARs to understand economic fluctuations. Surveys by Watson (1994), Christiano, Eichenbaum, and Evans (1997b), and Leeper, Sims, and Zha (1996) provide a fuller description of this literature.

reaction to the state of the economy, ϵ_t is an exogenous shock to monetary policy with unit variance, and σ is a scale parameter. The policy reaction function f incorporates the authority's preferences regarding counterstabilization actions, inflation-fighting activity, and so on. The residual ϵ reflects random, nonsystematic factors that affect policy decisions, such as political factors and the personalities, views, and composition of the Federal Open Market Committee.

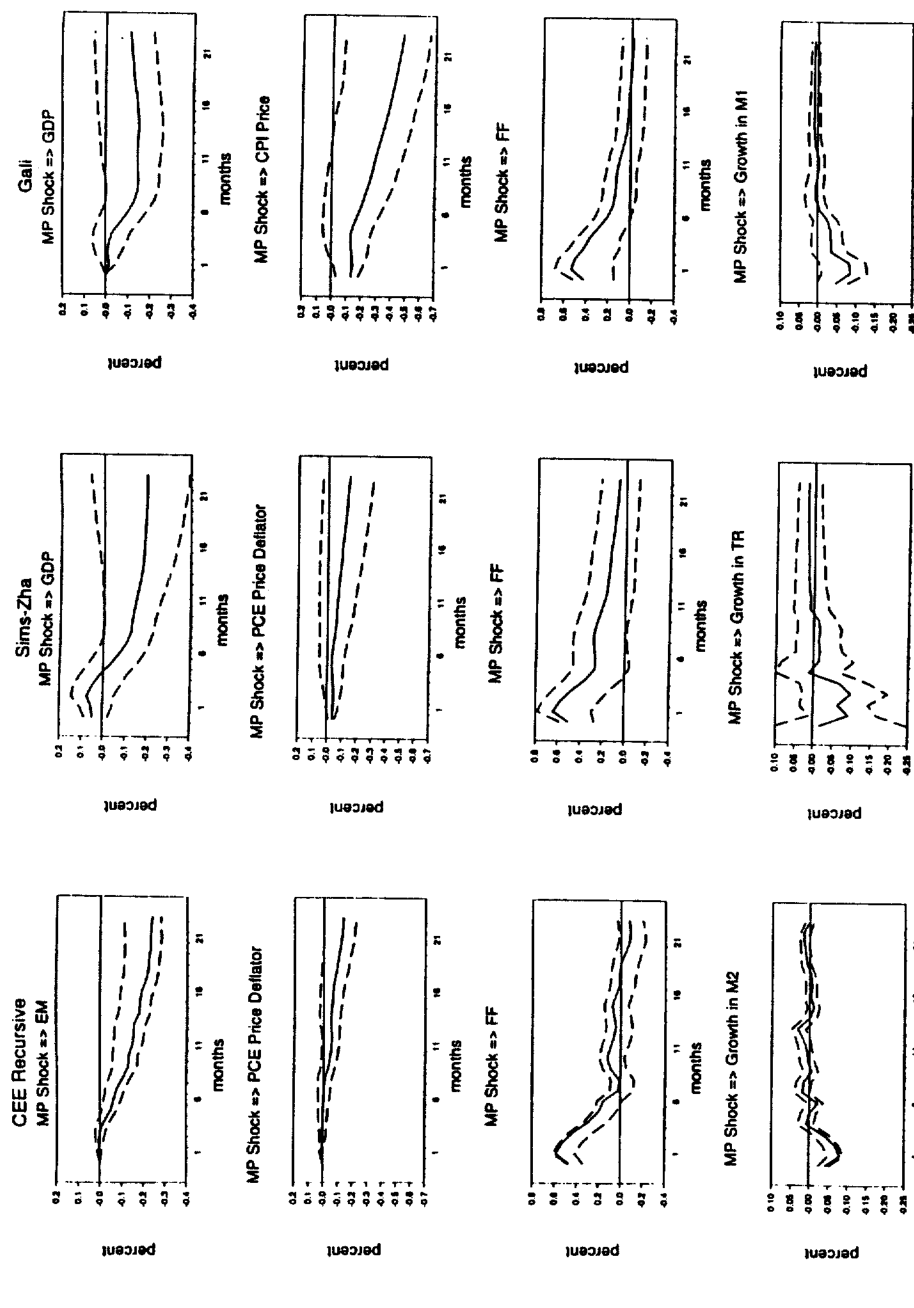
We will be considering in detail the dynamic effects of monetary-policy shocks on bond yields of various maturities. Specifically, let Z_t be a vector of macroeconomic variables at time t. In all cases we consider, FF_t is an element of Z_t. Let R_t^j denote a bond yield of maturity j months. The monetary policy rule (1) is estimated as one equation within a restricted version of the following structural VAR:

$$\begin{bmatrix} a & b \\ c & 1 \end{bmatrix} \begin{bmatrix} Z_t \\ R_t^j \end{bmatrix} = \begin{bmatrix} A(L) & B(L) \\ C(L) & D(L) \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ R_{t-1}^j \end{bmatrix} + \tilde{\sigma} \begin{bmatrix} \epsilon_t^Z \\ \epsilon_t^j \end{bmatrix} \tag{2}$$

where a is a square matrix with ones on the diagonal; b is a scalar; c is a row vector; $A(L)$ is a matrix polynomial in the lag operator L; $C(L)$ is a row vector polynomial; and $B(L)$ and $D(L)$ are scalar polynomials. The process $[\epsilon_t^Z \epsilon_t^j]'$ is an i.i.d. vector of mutually and serially uncorrelated structural shocks whose variance is the identity matrix; and $\tilde{\sigma}$ is a diagonal matrix. In this context, Ω_t consists of Z_{t-s} for $s \geq 1$ and certain contemporaneous values of Z_t depending upon the nonzero elements of a. Throughout our analysis, we maintain the assumptions that $b = 0$ and $B(L) = 0$. That is, neither contemporaneous nor lagged values of the bond yield enter the other equations in the system. These assumptions ensure that the shocks ϵ_t^Z are invariant to bond maturity j. The original papers that developed the approaches we use omitted bond yield data entirely, so these assumptions are consistent with their specifications. Our empirical results reported below are not appreciably different if we allow $B(L) \neq 0$. Bernanke, Gertler, and Watson's (1997) analysis maintained this latter assumption in a recursive identification scheme, and our related results are similar.[5] For an analysis that allows $b \neq 0$, see Gordon and Leeper (1994).[6]

[5]Bernanke, Gertler and Watson (1997) report impulse response functions from a VAR which includes both the 3-month and 10-year Treasury rates. They also maintain the assumption that $b = 0$.

[6]Leeper, Sims, and Zha (1996) discuss potential pitfalls in identifying monetary-policy shocks when contemporaneous bond-market data enter the monetary authority's information set and the expectations hypothesis is approximately correct. Bernanke and Woodford (1997) discuss potential indeterminacy problems for output and inflation when asset price data are employed in the policy rule.


Gali
MP Shock => GDP
MP Shock => CPI Price
MP Shock => FF
MP Shock => Growth in M1
Sims-Zha
MP Shock => GDP
MP Shock => PCE Price Deflator
MP Shock => FF
MP Shock => Growth in TR
CEE Recursive
MP Shock => EM
MP Shock => PCE Price Deflator
MP Shock => FF
MP Shock => Growth in M2
percent
months

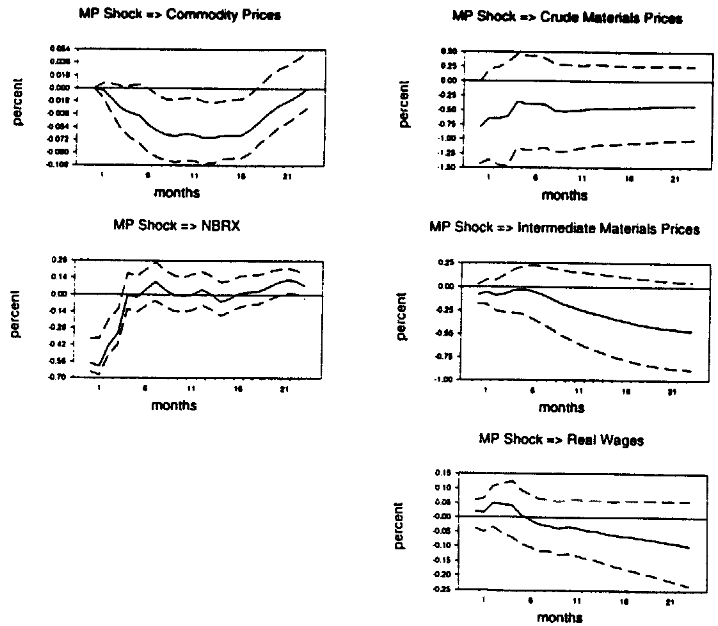


Figure 1

Responses of Macroeconomic Variables to Monetary-Policy Shocks

Notes to Figure 1: For each of the three identification strategies described in Section 2, this figure displays the responses of variables in the VAR system (other than bond yields) to a one-standard deviation contractionary monetary-policy shock. The first column displays impulse responses for the Christiano, Eichenbaum, and Evans recursive identification, described in Section 2.2. The responses reported are; employment (EM), personal consumption expenditure deflator (PCE Price Deflator), Federal funds rate (FF), growth in M2 money, an index of commodity prices, and the ratio of nonborrowed reserves to total reserves (NBRX). The second column displays impulse responses for the Sims-Zha nonrecursive identification, described in Section 2.3. The responses reported are; GDP, personal consumption expenditure deflator (PCE Price Deflator), Federal funds rate (FF), growth in total reserves (TR), an index of crude material prices, an index of intermediate material prices, and real wages. The third column displays impulse responses for the Gali identification, described in Section 2.4. The responses reported are: GDP, consumer price index (CPI), Federal funds rate (FF), and growth in M1 money. For all variables except the Federal funds rate, the responses are in percentage deviations from the steady state. For the Federal funds rate, the units are percentage points per annum. The solid lines plot the point estimates for the impulse responses; dashed lines give 95 % confidence intervals, as described in Sections 2.2 - 2.4.

2.2 *Christiano-Eichenbaum-Evans' recursive identification strategy*

Christiano, Eichenbaum, and Evans (1996b) assume that the monetary authority sees current prices and a measure of current economic activity when setting the monetary policy instrument. However, these prices and activity measures only respond to monetary policy with a one-month lag. Christiano, Eichenbaum, and Evans (1996b) employ monthly data. We have modified the data vector only slightly in order to facilitate comparisons with the theoretical model in Section 4. Specifically, the data vector is given by $Z^{CEE} \equiv (EM, P, PCOM, FF, NBR/TR, \Delta M2)'$, where: EM denotes the logarithm of nonagricultural payroll employment; P denotes the logarithm of the personal consumption expenditures deflator in chain-weighted 1992 dollars; $PCOM$ denotes the smoothed change in an index of sensitive materials prices; FF denotes the Federal funds rate; NBR/TR denotes the ratio of nonborrowed reserves plus extended credit to total reserves; and $\Delta M2$ denotes the log growth rate of the monetary aggregate M2.[7] As Christiano, Eichenbaum, and Evans (1996a) discuss, the inclusion of commodity prices in a recursively identified VAR mitigates anomalous responses of the price level from monetary-policy shocks (the "price puzzle" described by Sims (1992) and Eichenbaum (1992)).

The Christiano, Eichenbaum, and Evans monetary policy reaction function includes the contemporaneous values of EM, P, and $PCOM$ in the information set Ω_t in equation (1). Specifically,

$$FF_t = A_4(L)Z_{t-1}^{CEE} - a_{41}EM_t - a_{42}P_t - a_{43}PCOM_t + \tilde{\sigma}_{44}\epsilon_t^{CEE} \qquad (3)$$

where $A_4(L)$ is the fourth row of the matrix polynomial $A(L)$, and a_{ij} denotes the $(i,j)^{th}$ element of the matrix a. The monetary-policy shock ϵ_t^{CEE} is the fourth element of ϵ_t^Z and is assumed to be orthogonal to all other right-hand-side variables. This identification strategy has two important properties: (1) EM, P, and $PCOM$ do not respond contemporaneously to the monetary-policy shock, and (2) all of the other variables in the Z_t^{CEE} respond contemporaneously to the monetary-policy shock. In this sense, the identification of the monetary-policy shock ϵ_t^{CEE} is recursive.

The VARs were estimated over the sample period 1965:1 to 1995:12. Twelve lagged values were estimated in each equation, with the initial lags beginning in 1964:1. Column one in Figure 1 displays the impulse response functions for the Christiano, Eichenbaum, and Evans (*CEE Recursive*) monetary-policy shocks. Monte Carlo bootstrap methods were used to compute 95% confidence bands. The confidence bands are displayed around the point estimates of the impulse response functions, leading to generally asymmetric error bands (as suggested by Sims and Zha 1995b).

[7]For all our VARs, the logged data are also multiplied by 100 so that the impulse responses can be interpreted as percent deviations in all of our figures.

A one-standard deviation, contractionary monetary-policy shock leads to a 46-basis-point increase in the Federal funds rate on impact. The funds rate rises to its maximum of 58 basis points in the second month before falling thereafter. The funds rate response, therefore, is persistent but transitory: this response pattern holds for each of the identifications we consider. The other variables' responses seem consistent with most economists' prior expectations for a monetary-policy shock. The increase in the Federal funds rate occurs simultaneously with reductions in nonborrowed reserves relative to total reserves as well as M2. Employment and the PCE deflator are unchanged for several periods before falling persistently. Employment begins to fall before the PCE deflator, while commodity prices fall almost from the outset. Finally, the PCE deflator's response is *negligibly* positive for about six months. Of the three monetary policy identifications we consider, this is the largest price puzzle in Figure 1.

2.3 *Sims-Zha nonrecursive identification strategy*

Sims and Zha (1995a) criticize certain aspects of recursively identified VARs, such as in Christiano, Eichenbaum, and Evans. First, they note that commodity prices are set in auction markets, so commodity prices should respond immediately to innovations in monetary policy. (Recall that Christiano, Eichenbaum, and Evans assume that these prices respond only with a one-month lag.) Therefore, Sims and Zha advocate abandoning the recursiveness assumption of Christiano, Eichenbaum, and Evans in favor of a nonrecursive model: commodity prices should respond contemporaneously to monetary policy (as in Christiano, Eichenbaum, and Evans), but monetary policy in turn should also affect commodity prices contemporaneously. Secondly, Sims and Zha argue that measures of the price level and of real economic activity are compiled from survey data. These surveys take time to compile and are available to the monetary authority only with a delay. Consequently, Sims and Zha argue against treating these measures as contemporaneous inputs to policy.

Sims and Zha's empirical analysis used quarterly data. We use monthly analogues to their quarterly data series. Specifically, our data vector is given by $Z^{SZ} \equiv (Pcm, \Delta TR, FF, Pim, P, W, Y)'$, where: Pcm denotes the logarithm of crude materials prices; ΔTR denotes the log growth rate of total reserves; FF denotes the Federal funds rate; Pim denotes the log of intermediate goods prices; P denotes the log of the PCE deflator; W denotes the log of the real wage; and Y denotes the log of real GDP.[8] When Sims and Zha

[8]Our real GDP data are interpolated (from Leeper, Sims and Zha (1996)). Sims and Zha (1995a) used the quarterly GDP implicit deflator; our use of the PCE deflator is consistent with their choice of a price index with time-varying commodity bundle weights.

use the Federal funds rate as the monetary policy instrument, they select total reserves to be the monetary aggregate in the analysis. Our analysis follows their variable selection.[9]

The Sims and Zha monetary-policy reaction function includes the contemporaneous values of Pcm and ΔTR in Ω_t in equation (1). Specifically,

$$FF_t = A_3(L)Z_{t-1}^{SZ} - a_{31}Pcm_t - a_{32}\Delta TR_t + \tilde{\sigma}_{33}\epsilon_t^{SZ}. \qquad (4)$$

Where $A_3(L)$ is the third row of the matrix polynomial $A(L)$, and a_{ij} denotes the $(i,j)^{th}$ element of the matrix a. The monetary-policy shock ϵ_t^{SZ} is the third element of ϵ_t^Z. The Sims-Zha strategy is nonrecursive, because ϵ_t^{SZ} is allowed to be correlated with Pcm_t and ΔTR_t. (That is, Pcm_t and ΔTR_t are allowed to respond contemporaneously to a monetary-policy shock.) This correlation implies that ϵ_t^{SZ} cannot be recovered as the residual from an OLS regression. Furthermore, Sims and Zha's system of equations does not possess any predetermined variables which can be used as instruments for Pcm_t and ΔTR_t in equation (4). This leads Sims and Zha to full-information estimation methods. With these assumptions, the Sims and Zha strategy overcomes the two criticisms of the recursive identification scheme. Of course, the potential cost is in misspecifying the other equations in the full-information strategy.

Describing Sims and Zha's identification requires returning to the larger system in equation (2). Consider the system without reference to R^j, which is without loss of generality given our assumptions that $b = 0$ and $B(L) = 0$:

$$aZ_t = A(L)Z_{t-1} + \tilde{\sigma}_Z\epsilon_t^Z \qquad (5)$$

where $\tilde{\sigma}_Z$ denotes the upper-left square submatrix of $\tilde{\sigma}$ conformable with ϵ_t^Z. This system can be written as a reduced-form VAR:

$$[I - a^{-1}A(L)L]Z_t = a^{-1}\tilde{\sigma}_Z\epsilon_t^Z \equiv u_t \qquad (6)$$

Sims and Zha discuss identification as restrictions on a matrix G defined by $G = \tilde{\sigma}_Z^{-1}a$, which is natural given that $\epsilon_t^Z = Gu_t$. To achieve identification of their monetary-policy shock ϵ_{MP}, we follow Sims and Zha relatively closely

In all nonrecursive cases we considered, using the PCE deflator resulted in fewer price puzzles than Leeper, Sims and Zha's preferred use of the CPI. Also, we depart from Sims and Zha by excluding personal bankruptcies from the VAR.

[9] We also considered VARs with $\Delta M2$ in place of ΔTR, and the results were similar to the responses reported in Figure 1.

in specifying the G matrix:

$$\epsilon_t^Z \equiv \begin{bmatrix} \epsilon_{Pcm} \\ \epsilon_{MD} \\ \epsilon_{MP} \\ \epsilon_{Pim} \\ \epsilon_P \\ \epsilon_{w/p} \\ \epsilon_y \end{bmatrix} = \begin{bmatrix} G_{11} & G_{12} & G_{13} & G_{14} & G_{15} & G_{16} & G_{17} \\ 0 & G_{22} & G_{23} & 0 & G_{25} & 0 & G_{27} \\ G_{31} & G_{32} & G_{33} & 0 & 0 & 0 & 0 \\ G_{41} & 0 & 0 & G_{44} & G_{45} & G_{46} & G_{47} \\ G_{51} & 0 & 0 & 0 & G_{55} & G_{56} & G_{57} \\ G_{61} & 0 & 0 & 0 & 0 & G_{66} & G_{67} \\ G_{71} & 0 & 0 & 0 & 0 & 0 & G_{77} \end{bmatrix} \begin{bmatrix} u_{Pcm} \\ u_{TR} \\ u_{FF} \\ u_{Pim} \\ u_P \\ u_{w/p} \\ u_y \end{bmatrix}. \quad (7)$$

The first row indicates that Pcm is an *information variable*, responding to all structural shocks ϵ in the economy (other than the yield shock ϵ_t^j in equation (2)). The second row is a money-demand relationship. Our estimation constrains the coefficients on u_{TR} and u_{FF} to have the same sign, while u_y and u_P have opposite signs from u_{TR}. These sign restrictions ensure that the interest elasticity of money demand is negative, the output elasticity of money demand is positive, and that the price elasticity of demand for nominal balances is positive. The third row is the monetary policy reaction function. Rows four through seven indicate that Pim, P, w, and Y respond to the monetary-policy shock on impact *only indirectly through the effect of monetary policy on Pcm.*

The VARs were estimated over the sample period 1964:7 to 1995:12. Six lagged values were estimated in each equation, with the initial lags beginning in 1964:1. Column two in Figure 1 displays the impulse response functions for the Sims and Zha monetary-policy shocks. Following Sims and Zha (1995b), Bayesian Monte Carlo methods were used to compute 95% confidence bands. The confidence bands are displayed around the point estimates of the impulse response functions, leading to generally asymmetric error bands.

A one-standard deviation, contractionary monetary-policy shock leads to a 50-basis-point increase in the funds rate on impact, rising to 64 basis points in the second period. As with the Christiano, Eichenbaum, and Evans case, the funds rate falls thereafter. Total reserves fall over this period, although the initial response is close to zero. Prices fall on impact, with the response larger for crude materials' prices and smallest for the PCE deflator. Real GDP and wages display a small rise for the first five months before falling; however, the error bands for these impulses are wide. Broadly speaking, these responses are qualitatively similar to the recursive results, although they are estimated with less precision.

2.4 *Gali identification strategy using long-run restrictions*

Gali (1992) uses an alternative identification strategy which imposes a mixture of long-run restrictions and contemporaneous impact restrictions to identify four economic shocks. These four shocks are: an aggregate supply shock (which can be thought of as a technology shock); an aggregate demand, or "IS" shock; a money-demand shock, and a monetary-policy shock. A key identifying restriction (following Blanchard and Quah (1989)) is that only the aggregate supply shock can have a permanent effect on output. Gali's monetary-policy shock has no contemporaneous impact on output (like Christiano, Eichenbaum, and Evans but not Sims and Zha), but has a contemporaneous effect on prices (like Sims and Zha but not Christiano, Eichenbaum, and Evans).

We follow Gali in considering a four-variable autoregression. Gali's empirical analysis used quarterly data, while our analysis uses monthly data. The data vector is given by $Z^G \equiv (\Delta Y, FF, FF - \Delta P, \Delta M - \Delta P)'$, where: ΔY denotes the log difference of GDP; FF denotes the Federal funds rate; $FF - \Delta P$ denotes the real interest rate where ΔP is the log difference of the CPI; and $\Delta M - \Delta P$ denotes real M1 balances. To maintain comparability with the other procedures, we use the level of the Federal funds rate. (Gali used the first difference of the interest rate.) Since Gali's data are quarterly, we use the Leeper-Sims-Zha monthly data set for monthly GDP, Fed funds rate, CPI, and M1.[10] Gali's monetary policy reaction function can be represented as

$$FF_t = A_2(L)Z^G_{t-1} - a_{21}\Delta Y_t$$

$$-a_{23}(FF_t - \Delta P_t) - a_{24}(\Delta M_t - \Delta P_t) + \tilde{\sigma}_{22}\epsilon^G_t \quad (8)$$

where $A_2(L)$ is the second row of the matrix polynomial $A(L)$, and a_{ij} denotes the $(i,j)^{th}$ element of the matrix a. The monetary-policy shock ϵ^G_t is the second element of ϵ^Z_t, and is potentially correlated with time t explanatory variables (as in the Sims-Zha system). Identification is achieved with six restrictions on the covariance structure of the innovations. First, the monetary-policy, money-demand and IS shocks have no long-run effects on output; these restrictions identify the supply shock. Second, the monetary-policy and money-demand shocks have no contemporaneous effect on output; knowledge of the supply shock and these two restrictions identifies the IS shock. Third, one additional identifying restriction is necessary to identify the remaining two shocks. One of the restrictions that Gali considers deletes the price data from the monetary authority's contemporaneous information set. In equation (8), this imposes the coefficient restriction that $a_{23} = -a_{24}$, leaving only two contemporaneous coefficients to estimate with two available

[10] We follow Gali in using the CPI rather than the PCE deflator for two reasons. First, Gali conducted several sets of unit root and cointegration tests in order to justify his data transformations. Consequently, we used his P and $M1$ to maintain comparability. Second, the CPI data system delivered more plausible impulse-response functions for most of the shocks than the PCE deflator data system. Since the CPI does not get revised, its stochastic trend properties may be more consistent with Gali's unit root assumptions.

instruments (the supply and IS shocks). This identifies the monetary-policy shock.[11]

The VARs were estimated over the sample period 1964:8 to 1995:12. Six lagged values were estimated in each equation, with the initial lags beginning in 1964:2.[12] Column three in Figure 1 displays the impulse response functions for the Gali monetary-policy shocks. Monte Carlo bootstrap methods were used to compute 95% confidence bands. The confidence bands are displayed around the point estimates of the impulse response functions.

A one-standard deviation, contractionary monetary-policy shock increases the Federal funds rate on impact by 41 basis points. In the second period it increases by 53 basis points, and falls thereafter. M1 growth falls during this period, indicating a liquidity effect. As with the Sims and Zha policy shock, the price level falls on impact and declines further after about six months. As with the Christiano, Eichenbaum, and Evans policy shock, real activity (as measured by monthly real GDP) is about flat for four months and then falls.

3 The response of bond yields to exogenous monetary-policy shocks

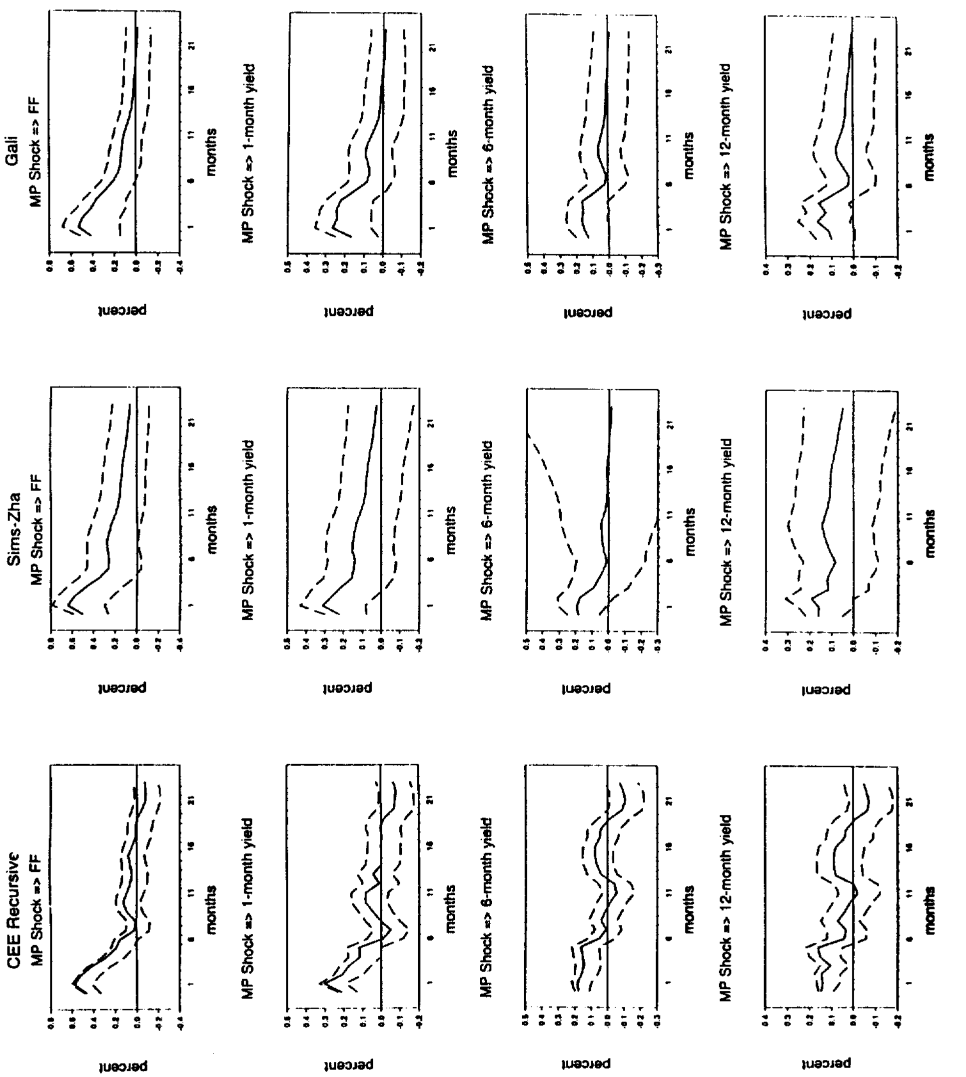
3.1 *Impulse response*

Figure 2 plots the estimated responses of bond yields to a one-standard-deviation contractionary monetary-policy shock. Bond yields are measured as continuously-compounded annualized returns on zero-coupon bonds. The yields from 1959:01-1991:02 are monthly data taken from McCulloch and Kwon (1993). For the period 1991:03-1995:12, we use yields computed by Robert Bliss using the McCulloch/Kwon procedure. (See Bliss (1994)).[13] The solid lines give the point estimates of the impulse responses; the upper and lower dashed lines give the boundaries of the 95% confidence region. The

[11] Gali (1992) alternatively considers: (1) deleting only output from the contemporaneous information set of the monetary authority, and (2) explicitly imposing a homogeneity restriction on the money-demand equation in his structural VAR. Gali reported that his results were largely robust across these alternative identification restrictions, and our implementation of these restrictions also produced qualitatively similar results.

[12] "Double-differencing" to impose the long-run restrictions in identifying the supply shock uses up the additional lag beginning in 1964:1.

[13] McCulloch and Kwon's (1993) data on zero-coupon bond yields are derived from a tax-adjusted cubic spline discount function, as described in McCulloch (1990). A more detailed explanation can be found in McCulloch and Kwon (1993). Unlike McCulloch and Kwon (1993), Bliss (1994) does not tax-adjust the bond yields. However, under the current tax code, the requisite tax adjustment in the McCulloch-Kwon procedure is negligible. From 1987:01 through 1991:02 (the last date where we have an overlap between the two data sets), the McCulloch/Kwon data and the Bliss data are virtually identical.



CEE Recursive
Sims-Zha
Gali
MP Shock => FF
MP Shock => 1-month yield
MP Shock => 6-month yield
MP Shock => 12-month yield
percent
months

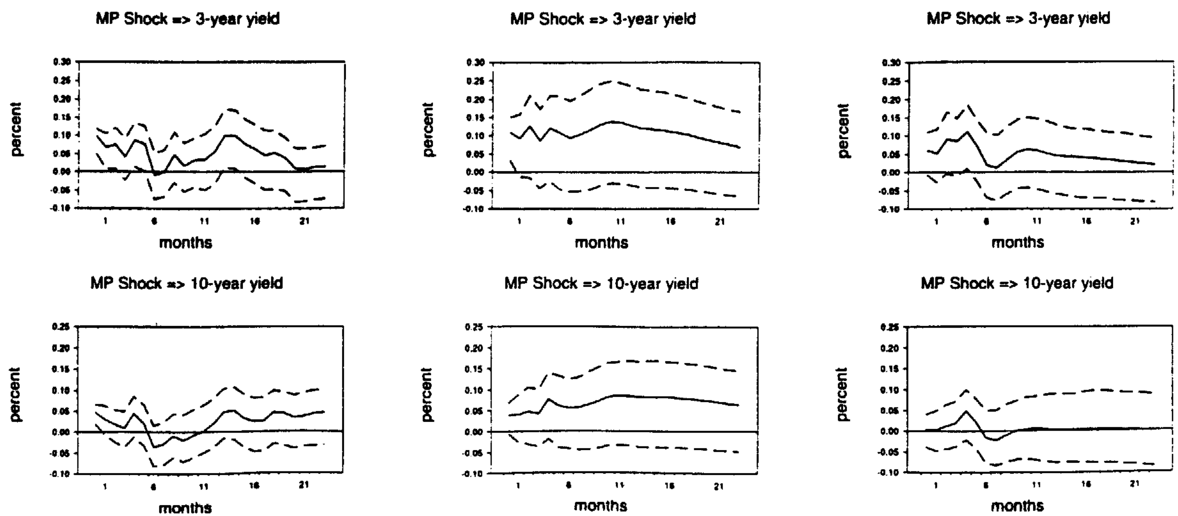


Figure 2

Responses of Bond Yields to Monetary-Policy Shocks

Notes to Figure 2: For each of the three identification strategies described in Section 2, this figure displays the response to a one-standard deviation contractionary monetary-policy shock of the Federal funds rate (first row) and of the continuously-compounded yields for zero coupon bonds of maturities one month, six months, one year, three years, and ten years (rows two through six). The first column displays impulse responses implied by CEE recursive identification (described in Section 2.2); The second column displays impulse responses implied by the Sims-Zha nonrecursive identification (described in Section 2.3); The third column displays impulse responses implied by the Gali identification (described in Section 2.4). For all these impulse responses, the units are percentage points per annum. The solid lines plot the point estimates for the impulse responses; dashed lines give 95% confidence intervals, as described in Sections 2.2 - 2.4.

plots trace the responses over 24 months. Each of these responses is measured in percent deviation from the nonstochastic steady state. We display the responses for bond maturities of one month, six months, one year, three years, and ten years. According to all three of the identification strategies, the policy shock increases the one-month rate by approximately 20 basis points in the period when the shock occurs. This response is statistically significant in each case. The one-month rate continues to climb in the following months, and then falls rapidly, with the effect of the shock dissipating after a year. The six-month and the twelve-month rates display qualitatively similar patterns, although the magnitude of the response decreases for the longer-term bonds. When we move to even longer-maturity bonds, the initial effect diminishes substantially as maturity increases: The initial response of the three- year bond is only around 9 basis points, falling to less than 5 basis points for the ten-year bonds. The main qualitative discrepancy among the three identification strategies is that the bond-yield responses die off somewhat more slowly in the Sims and Zha identification than in either the Christiano, Eichenbaum, and Evans or the Gali identifications. Interestingly, these results are roughly comparable to Cook and Hahn's (1989) estimates of the effects on interest rates after a publicly announced change in the Federal funds rate. They find that in response to a 100-basis-point increase, short rates rise about 50 basis points, while long rates rise about 10 basis points.

Summarizing these results, there is a large and significant but relatively short-lived effect on short rates, with a decreasing effect on longer maturities. In other words, there is not a parallel upwards shift of the term structure in response to these monetary-policy shocks; rather, the shock causes the yield curve to flatten.[14] An alternative way to portray these patterns is to look at the effect of a monetary shock on the shape of the yield curve. One way to summarize this shape is to take a quadratic approximation of the yield curve at each date. We do so by regressing all interest rates at a given date on a constant, maturity, and squared maturity, and treating the parameter estimates (denoted *intercept, slope*, and *curvature*, respectively) as the coefficients of this quadratic approximation. (Note that these coefficients are time-varying, since the regression only involves interest rates at a given date, and is reestimated each month.) To portray the way the shape of the

[14] As a check on subsample robustness, we replicated the analysis displayed in Figures 1 and 2 using data from 1983:1 through 1995:12. The broad patterns characterizing these figures also appear in this subsample analysis. In particular, the response of yields to a monetary contraction are rather short-lived, decreasing with the maturity of the yield. Not surprisingly, the impulse-response functions are estimated with less precision. The standard deviation of the monetary-policy shock in all identifications is smaller than with the full sample, ranging from 15 to 25 basis points. The recursive identification displays a price puzzle, although the error bands are wide. The other identifications imply essentially flat price responses.

yield curve responds to a monetary shock, we estimate VARs, analogous to those described above, in which the interest rate is replaced by one of these three time-varying coefficients.

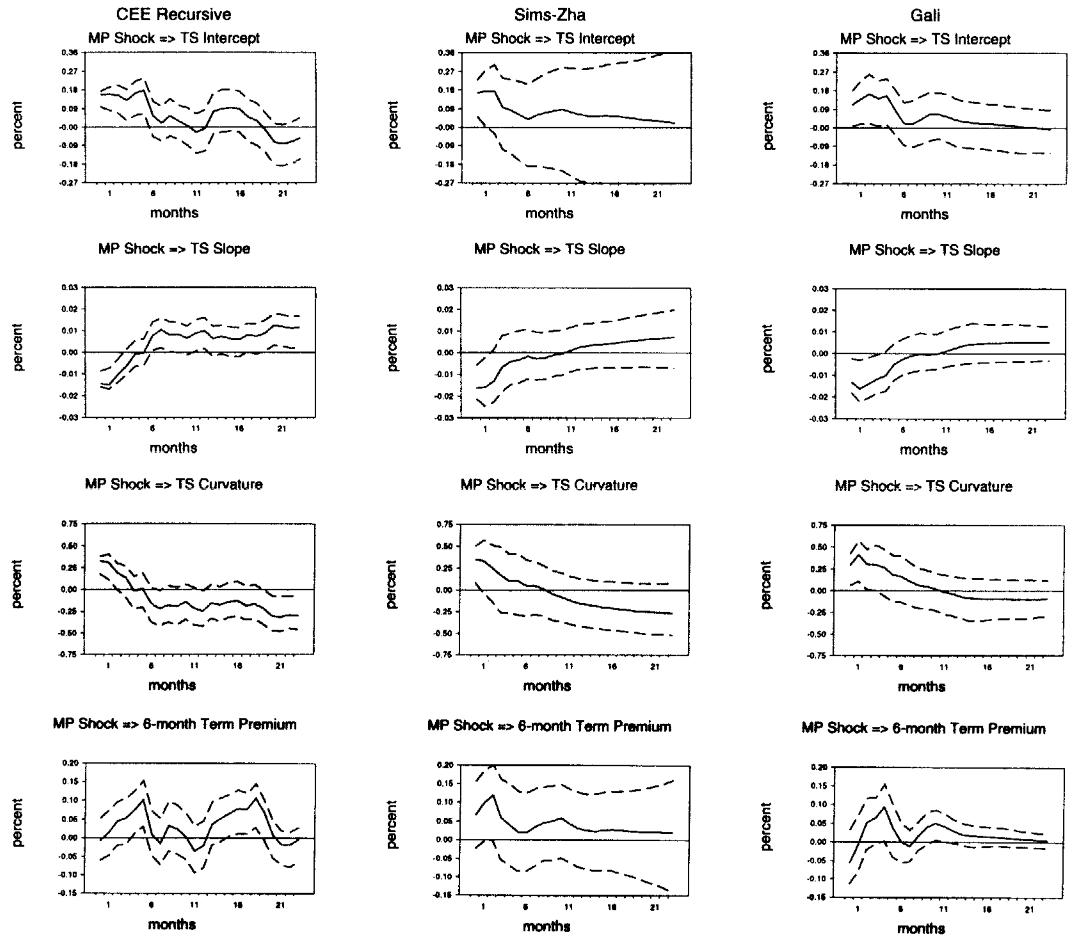
The resulting impulse responses are displayed in Figure 3. A monetary shock raises the level of the yield curve, decreases the slope, and reduces the curvature. (The positive response of *curvature* denotes a reduction in curvature because the average yield curve is concave, so the average value of *curvature* is negative.) The positive response of *intercept* looks very much like the response of the one-month interest rate. For all three identifications, this response dissipates within six months. The effects on *slope* and *curvature* also dissipate within four to six months.

3.2 *Variance decompositions*

The impulse responses suggest that monetary policy is an important determinant of short-run interest-rate variability, at least for the shorter-term rates. To study this question directly, consider the variance decompositions displayed in Table 1. The table gives the point estimates of the fraction of the one-month-ahead, six-month-ahead, and 24-month-ahead conditional variance of five bond yields attributable to the monetary-policy shock, as identified by each of the three identification strategies. According to this table, monetary-policy shocks account for 17% - 18% of the conditional variance of the one-month interest rate. The fraction of the six-month-ahead variance accounted for by the monetary-policy shock decreases sharply with maturity. It is still nontrivial for the one-year interest rate (9% - 11%), but rapidly becomes negligible as maturity lengthens.

The 24-month-ahead conditional variance can be interpreted as a proxy for the unconditional interest-rate variance. According to the Christiano, Eichenbaum, and Evans and Gali identifications, monetary-policy shocks account for a relatively smaller fraction of the long-run variance of interest rates (around 7% for the one-month rate, and less for the longer-term rates). This reflects the rapid decay in the impulse responses implied by these identification strategies. The Sims and Zha identification attributes somewhat more of the long-run variance to monetary-policy shocks, due to the greater persistence of the impulse responses implied by that identification strategy.

The impulse responses and variance decompositions suggest that the monetary-policy shock resembles the "slope" factor identified in the finance literature. In particular, Litterman and Scheinkman (1991) and Dai and Singleton (1997) estimate factor models of the term structure in which the three factors shift the level, slope, and curvature of the yield curve, respectively. Litterman and Scheinkman (1991) find that the level factor accounts for about 90% of the unconditional variability of yields across the maturity


CEE Recursive
Sims-Zha
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MP Shock => TS Intercept
MP Shock => TS Slope
MP Shock => TS Curvature
MP Shock => 6-month Term Premium
percent
months

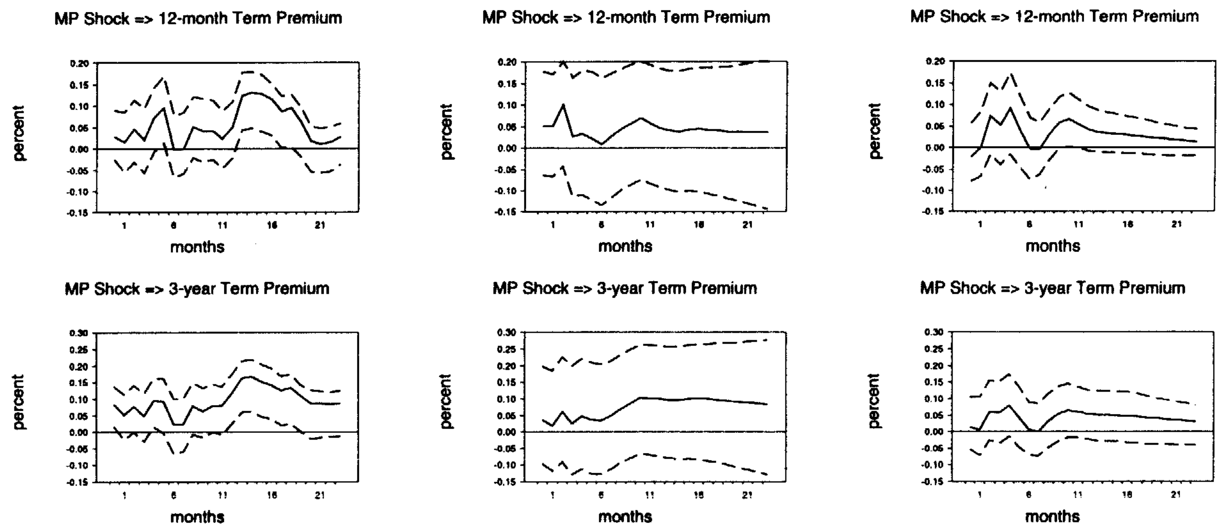


Figure 3

Responses of Term Structure Descriptors and Term Premiums to Monetary-Policy Shocks

Notes to Figure 3: The interpretation of Figure 3 is identical to that of Figure 2, except that, instead of displaying responses of bond yields to a monetary-policy shock, the first three rows display responses of the three term structure descriptors *intercept, slope* and *curvature*. These term structure descriptors are computed as described in Section 3.1. The last three rows display the responses of the j-period term premium, TP_t^j defined in equation (9).

Table 1:
Fraction of Yield Variance Explained by Monetary-Policy Shock

A. CEE Identification
Bond Maturity

Horizon	One month	Six months	One year	Three years	Ten years
1-month	16.7%	14.8%	11.0%	6.5%	2.9%
6-months	16.9%	12.9%	10.8%	4.8%	1.5%
24-months	6.9%	5.3%	4.6%	3.0%	1.4%

B. SZ Identification
Bond Maturity

Horizon	One month	Six months	One year	Three years	Ten years
1-month	13.6%	12.3%	10.2%	6.6%	1.5%
6-months	18.1%	6.9%	8.7%	7.6%	3.6%
24-months	15.4%	3.7%	11.3%	14.5%	10.3%

C. Gali Identification
Bond Maturity

Horizon	One month	Six months	One year	Three years	Ten years
1-month	7.2%	4.9%	3.5%	1.9%	0.0%
6-months	17.1%	13.7%	11.2%	5.0%	3.2%
24-months	7.4%	2.9%	2.9%	2.2%	0.2%

Notes to Table 1: The estimated fraction of the one-month-ahead, six-month-ahead, and 24-month-ahead conditional bond-yield variance attributable to monetary-policy shocks is displayed. Panel A displays results for the Christiano, Eichenbaum, and Evans recursive strategy for identifying monetary-policy shocks (described in Section 2.2); Panel B displays results for the Sims-Zha nonrecursive identification strategy (described in Section 2.3); and Panel C displays results for the Gali identification strategy that incorporates long-run restrictions (described in Section 2.4). For each identification strategy, results are displayed for bonds maturing in one month, six months, one year, three years, and ten years.

spectrum, with the slope factor accounting for most of the rest. However, Dai and Singleton (1997) note that the slope factor accounts for a good deal of the short-run variability of the short-term interest rate. The slope factor is less important for the unconditional variability because it has only a transitory impact, with a half-life of about 4 months. (In contrast, the level factor induces more persistent responses in the yield curve. The half-life of a level-factor impulse is estimated at approximately 4-1/2 years.) All of these characteristics correspond closely to our result for the monetary-policy shock. One might conjecture that the slope factor identified by Litterman and Scheinkman and Dai and Singleton is, in part, driven by monetary policy.

At the one-month horizon for all maturity yields, most of the forecast error variance is attributed to the yield's own shock (ϵ_t^j in equation (2)). Depending on the identification strategy, ϵ_t^j accounts for 73-75% of the one-month-ahead error variance for the one-month yield; 75-80% for the twelve-month yield; and 90-95% for the ten-year yield. The similarity across the three different VARs reflects a general lack of predictability of asset prices at the shortest horizons. At longer horizons, the error variances are substantially smaller and the range of variance estimates is wider. The ϵ_t^j shock accounts for 15-35% of the 24-month-ahead error variance for the one- month yield; 16-45% for the twelve-month yield; and 24-43% for the ten-year yield.[15] Finally, in each of our identification schemes, the yield shock is independent of the monetary- policy and other shocks. Consequently, the large error variances attributed to these yield shocks simply reflect financial market volatility that is unrelated to the structure of the economy. The size of these error variances provides no evidence for or against the identification strategies studied here.

3.3 *Term premiums*

Monetary-policy shocks could affect longer rates either through their effect on expected future short rates or by affecting term premiums. To distinguish between these two alternatives, recall that R_t^j denotes the j-period continuously-compounded bond yield. Let us define the j-period term premium, TP_t^j, by

$$TP_t^j \equiv R_t^j - \frac{1}{j}\sum_{i=0}^{j-1} E_t R_{t+i}^1 \tag{9}$$

That is, the term premium is the difference between the j-period interest rate and the average of expected future one-period interest rates over the next j periods. The expectations theory of the term structure is the hypothesis that term premiums are time-invariant. It can be shown (see Bekaert,

[15]These results are consistent with the greater predictability of long-horizon returns versus short-horizon returns, documented in Fama and French (1988).

Hodrick, and Marshall (1997a)) that the expectations hypothesis is equivalent to the hypothesis that the intertemporal marginal rate of substitution in nominal wealth is conditionally homoskedastic, in the strong sense that *all* conditional higher moments are time-invariant. Of course, the expectations theory has been rejected decisively in US data,[16] so, empirically, TP_t^j varies through time. It is of interest, therefore, to see whether monetary-policy shocks affect longer yields primarily through their effect on expected future short yields, or whether they directly affect term premiums. This is an important issue in its own right, and it may serve to indirectly inform us about the way monetary policy affects the elusive intertemporal marginal rate of substitution in wealth.

To help understand the sources of time-variation in term premiums, we compute the response of TP_t^j to the monetary-policy shock in our model as the difference between the contemporaneous response of R_t^j and the average of the first j-step responses of R_t^1. In Figure 3, we display these responses, along with the 95% confidence intervals, for j = six months through three years.[17] For all three identification strategies, the point estimates imply that the six-month term premium responds positively to a contractionary monetary-policy shock, with the maximal response of approximately 10 basis points occurring between two and five months after the initial impulse. However, these estimates are imprecisely estimated, leaving a zero response within the range of high probability. For the longer maturities, the Christiano, Eichenbaum, and Evans identification implies a rather long-lived term premium response. For example, the twelve-month term premium rises to 12 basis points after one year. However, these responses are not found in the other identification strategies. We conclude that the evidence for nonzero term premium responses is decidedly mixed.

3.4 *Real yields and expected inflation premiums*

According to Figures 1 and 2, nominal bond yields and inflation move in opposite directions in response to a monetary-policy shock, so the nominal yield response clearly represents a liquidity effect, rather than an expected inflation effect. To quantify the magnitude of this liquidity effect, we compute

[16] However, the expectations theory fares far better in other countries. For example, the expectations theory cannot be rejected using data from the UK, and the rejections are far less decisive with German data. See Bekaert, Hodrick, and Marshall (1997a), Hardouvelis (1994), Jorion and Mishkin (1991).

[17] To estimate the response of term premia, we must estimate a seven-variable VAR, including both the one-month interest rate and the j-month rate. It is problematical to perform this exercise for the longer-maturity interest rates, since it involves summing the first j responses of the long interest rate. In the case of the ten-year rate, for example, we would have to sum the first 120 responses. We have little confidence in the point estimates over this long a horizon.

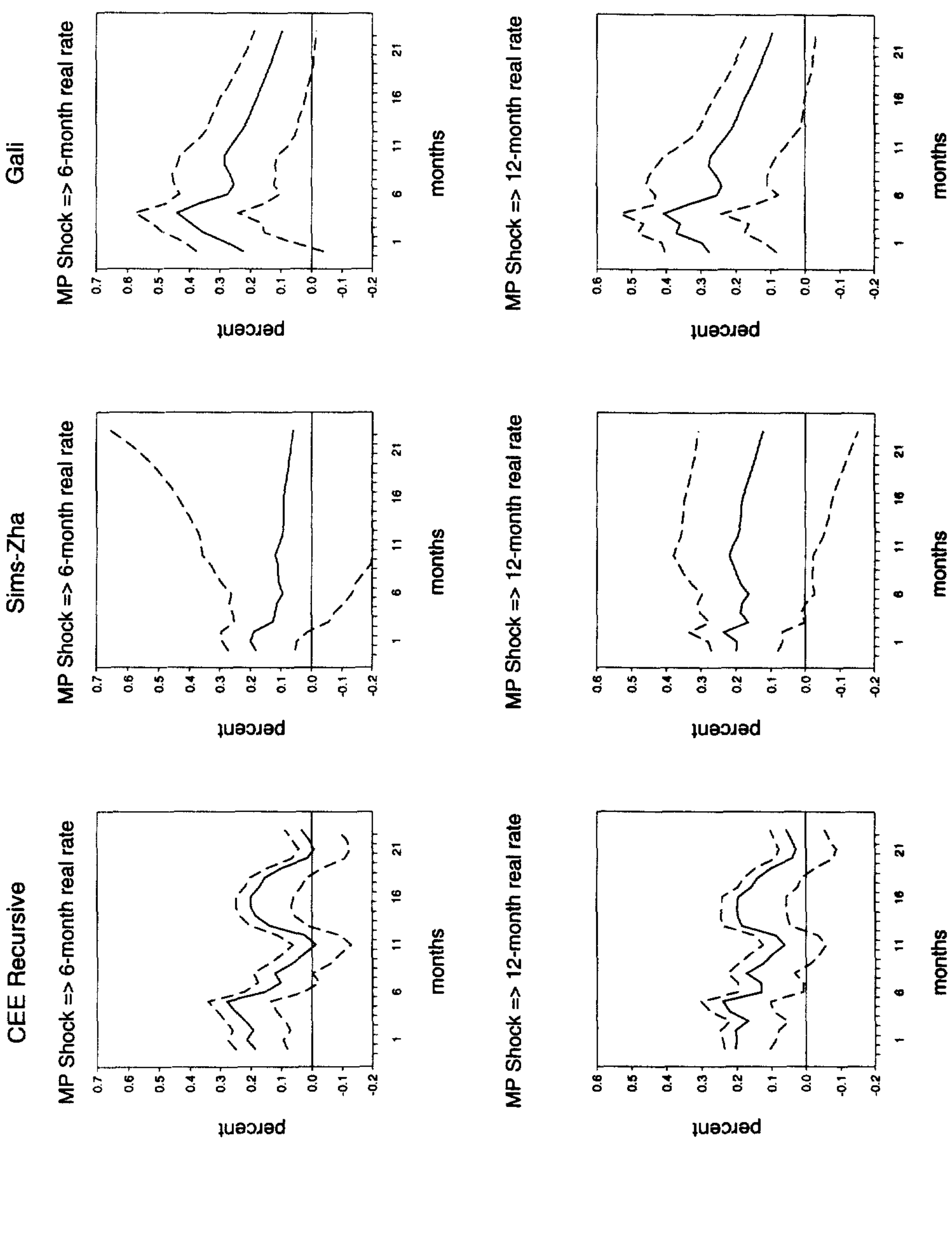
the response of the real j-month yield to the monetary-policy shock. We do so by subtracting from the i^{th} response of the j-month nominal yield the average (annualized) inflation response from step $i+1$ to step $i+j$. These computations are displayed in Figure 4 for j = six months through three years. According to all three identification strategies, real yields respond significantly to monetary-policy shocks for all maturities displayed. Notably, both the Christiano, Eichenbaum, and Evans and Gali identifications imply that the positive response of the twelve-month and three-year real yields is economically meaningful (about 20 basis points) and highly persistent, remaining positive with high probability for a year or more. (The point estimates from the Sims and Zha identification tell a similar story, although the error bands are much wider for this identification strategy.)

4 Can a dynamic equilibrium model replicate these patterns?

The empirical results for each of the three identification strategies are remarkably similar. A contractionary monetary shock causes a substantial rise in short-term nominal yields, with a progressively smaller response as the bond maturity is lengthened. This in turn flattens the slope and curvature of the yield curve. These responses are rather transient, fully dissipating between six months and one year. They represent pure liquidity effects: The responses of real yields are significant, and generally exceed the response of the nominal yields. There is some evidence that term premiums also respond positively, at least for the shorter yields. The effect of a contractionary monetary- policy shock on noninterest-rate variables is to induce a slow, persistent decline in real economic activity, and an even slower, more persistent fall in the price level.

What features must a dynamic equilibrium monetary model possess if it is to match these empirical patterns? First, we must have some sort of nominal rigidity. In a simple variant of the cash-in-advance paradigm, the interest-rate response to a monetary contraction reflects only a fall in expected inflation. This would induce a *decline* in nominal interest rates and (to the extent that inflation represents a distortionary tax) a *rise* in output, precisely the opposite of the observed patterns. Second, we need the output response to be stronger, at least in the short run, than the price response. If not, the expected inflation effect on interest rates would dominate the liquidity effect. (Furthermore, the impulse responses in Figure 1 indicate that the output response tends to precede the price response.) Finally, we need some mechanism to induce persistent responses to monetary impulses.

There are a number of dynamic models of monetary nonneutrality in the literature that potentially could satisfy these requirements, including sticky-price models (see, for examples, McCallum (1994b) and Goodfriend and King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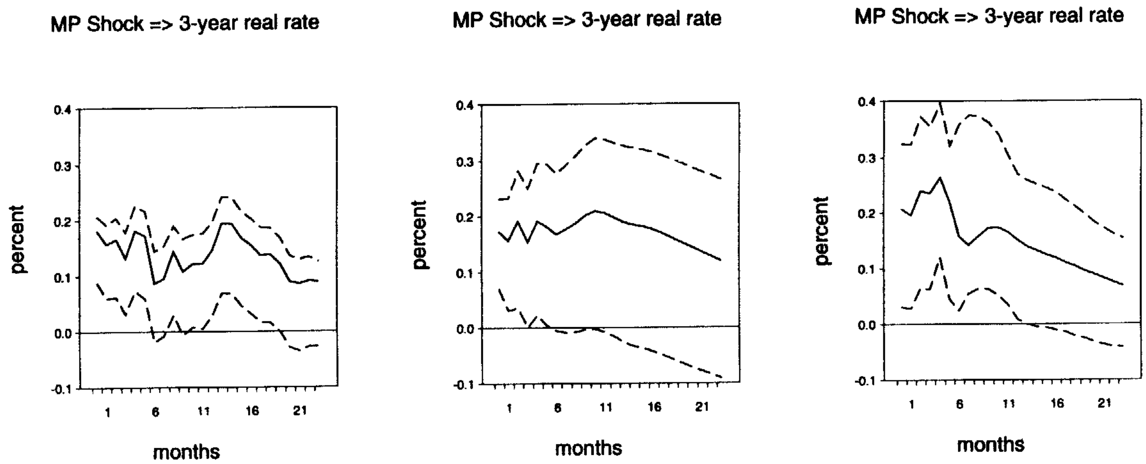


Figure 4

Responses of Real Interest Rates to Monetary-Policy Shocks

Notes to Figure 4: The interpretation of Figure 4 is identical to that of Figure 2, except that, instead of displaying responses of nominal bond yields to a monetary-policy shock, the responses of real bond yields of maturities six months, 12 months, and three years are displayed. The first column displays impulse responses for the CEE recursive identification, described in Section 2.2. The second column displays impulse responses for the Sims-Zha nonrecursive identification, described in Section 2.3. The third column displays impulse responses for the Gali identification, described in Section 2.4. In computing the real bond yields, the CEE identification (column 1) and the Sims-Zha identification (column 2) compute the inflation rate using the personal consumption expenditure deflator; the Gali identification (column 3) uses CPI inflation. In all cases, units are in percentage points per annum.

(1997)). Rather than exploring the implications of several different classes of monetary models, our goal is more limited. We ask whether a particular type of monetary model can replicate the empirical patterns documented in Section 3. The class of dynamic models we study incorporates the limited participation assumption introduced in Lucas (1990) and studied in Fuerst (1992) and Christiano and Eichenbaum (1995). The benchmark model we use is a variant of the model analyzed in Christiano, Eichenbaum, and Evans (1997c). The nominal rigidity in this class of models is that households must decide how much cash (denoted "Q_t") to use in the goods market *before* the monetary-policy shock is revealed. Furthermore, it is assumed that Q_t cannot be adjusted without cost. Rather, there is a cost in leisure time that must be paid whenever consumption money is changed from period to period. This adjustment cost allows real effects of a monetary-policy shock to propagate dynamically through time. The following is a brief description of the model. (See Christiano, Eichenbaum, and Evans (1997c) for a more detailed description.)

4.1 *Basic set-up*

There are three types of entities: households, firms, and a financial intermediary, plus a government whose sole function is to create money. The households own the firms and the financial intermediary, so all profits from these entities are paid to the households. Households' purchases of consumption and new capital are subject to a cash-in-advance constraint, to be described below.

Since the focus of this paper is on the effects of monetary policy, we assume that, unlike the empirical models of Section 2, monetary policy is the only source of randomness in this economy. Each period, the government injects a quantity of money X_t as a transfer to the financial intermediary. The total quantity of money in the economy evolves according to $M_{t+1} = M_t + X_t$. For convenience in calibrating the model to the impulse responses described above in Section 3, we assume that net money growth $x_t \equiv M_{t+1}/M_t - 1$ evolves as a stationary moving-average process:

$$x_t = x + \theta_0 \epsilon_t + \theta_1 \epsilon_{t-1} + \theta_2 \epsilon_{t-2} + \theta_3 \epsilon_{t-3} + \theta_4 \epsilon_{t-4} \tag{10}$$

where x denotes the nonstochastic steady-state rate of money growth, and $\{\epsilon_t\}$ is a sequence of i.i.d. standard normal shocks. Equation (10) is an exogenous monetary policy rule, while equation (1) is an endogenous monetary policy rule. Notice that the monetary authority could be using equation (1) to set the Federal funds rate and yet the Wold representation for money growth would be a moving average of current and lagged exogenous shocks, as approximated by the finite-order MA in equation (10). The observational equivalence of these rules allows us to estimate an endogenous monetary-policy rule from the data, and then use the equivalent exogenous policy rule as the driving process in the model.[18] The $\{\theta_i\}_{i=0}^4$ coefficients in equation (10) are computed directly as the impulse response coefficients from row four in Figure 1.[19] In all cases we truncate the MA process at an MA(4). As the confidence bands in Figure 1 show, this truncation seems reasonable.

4.2 *Households*

The representative household's choice variables in period t are: consumption C_t, labor L_t, cash set aside for purchases Q_t, money M_{t+1}, capital K_{t+1}, and a portfolio of zero-coupon bonds with maximum maturity of n periods, denoted $\{B_t^j\}_{j=1}^n$. In this notation B_t^j denotes a bond purchased at date t paying one dollar at the end of date $t+j-1$. The household takes as given the nominal rental rate on capital, r_t, the dollar price b_t^j of a bond maturing at the end of period $t+j-1$, as well as the dollar prices of labor and consumption goods, W_t and P_t.

The timing is as follows: at the beginning of period t, the household carries over from the end of the previous period M_t, K_t, and bonds of maturities 2 through $n, \{B_{t-1}^j\}_{j=2}^n$. (One-period bonds purchased in period $t-1$ pay off at the end of period $t-1$.) Before the monetary-policy shock in period t is revealed, the household must set aside Q_t dollars to finance purchases subject to the cash-in-advance constraint. The household takes its remaining financial assets (money-holdings $M_t - Q_t$ and holdings of zero-coupon bonds) to the financial intermediary. The monetary-policy shock in period t is then revealed. Having seen the shock, the household rebalances its portfolio by purchasing from the intermediary bonds of maturities 1 through n. The portfolio constraint facing the household in these transactions is:[20]

$$\sum_{j=1}^{n} b_t^j B_t^j \leq M_t - Q_t + \sum_{j=2}^{n} b_t^{j-1} B_{t-1}^j. \tag{11}$$

The household then rents its capital K_t at nominal rental rate r_t and sells its labor L_t to a firm for nominal wage W_t. Wages are paid in money that

[18]If the economy is subjected to multiple shocks, such as technology and preference shocks, and the monetary authority employs an endogenous monetary-policy rule like equation (1), then its exogenous representation would be a multivariate form of equation (10). Suppose we restrict our interests to only certainty equivalent solutions. In this setting, our analysis of impulse responses following a monetary-policy shock can be interpreted as a conditional analysis that holds other shocks fixed.

[19]Recall that the log money-growth data has been scaled up by a factor of 100.

[20]As long as $b_\tau^1 < 1$ (equivalently, the one-period net nominal interest rate is positive), the household sells its entire money holdings (net of Q_t) to the intermediary.

can be used immediately for purchases of consumption and new capital. The cash-in-advance constraint can therefore be written:

$$P_t(C_t + (K_{t+1} - (1-\delta)K_t)) \leq Q_t + W_tL_t, \tag{12}$$

where δ denotes the capital depreciation rate. Finally, at the end of the period, the firm pays out all profits to the household as a dividend D_t, and the financial intermediary redeems all maturing bonds B_t^1 and pays out all of its profits to the household as a dividend F_t. The flow budget constraint for nominal household wealth can therefore be written:

$$M_{t+1} \leq F_t + D_t + B_t^1 + r_tK_t + Q_t + W_tL_t - P_t(C_t + (K_{t+1} - (1-\delta)K_t)). \tag{13}$$

Note that, in equation (13), one-period bonds purchased in the period t bond market are assumed to pay off at the end of period t, rather than at the beginning of period $t+1$. This is without loss of generality, since the bond payoff is known with perfect certainty and cannot be used to offset the period t cash-in-advance constraint. (See equation (12).) We adopt this bookkeeping convention as a convenience: it ensures that all cash resides with the household at the end of each period.

Let H_t denote the time cost of adjusting Q. This adjustment cost is assumed to have the following form:

$$H_t = H\left(\frac{Q_t}{Q_{t-1}}\right) =$$

$$d\left\{exp\left[c\left(\frac{Q_t}{Q_{t-1}} - 1 - x\right)\right] + exp\left[-c\left(\frac{Q_t}{Q_{t-1}} - 1 - x\right)\right] - 2\right\}, \tag{14}$$

where x (the steady-state growth rate of money in equation (10)) is the net growth rate in Q_t in a nonstochastic steady state.

In period t, the household chooses $C_t, Q_t, L_t, M_{t+1}, K_{t+1}$, and $\{B_t^j\}_{j=1}^n$ to maximize

$$E_{-1}\sum_{t=0}^{\infty}\beta^t U(C_t, L_t, H_t), \tag{15}$$

subject to (11), (12), and (13), where

$$U(C,L,H) = \left[C - \psi_0\frac{(L+H)^{(1+\psi)}}{1+\psi}\right]^{(1-\gamma)} /(1-\gamma). \tag{16}$$

This utility function has the property that the income effect on leisure is zero. Everything else equal, this tends to magnify the output response from a monetary shock. Intuitively, the household's labor supply does not decrease when money expands. Parameter ψ is the inverse of the elasticity of labor supply. Parameter γ is a curvature parameter that affects the household's degree of risk aversion. Parameter ψ_0 is purely a scaling parameter. Households make all date t choices except one as functions of information known at date t and earlier. The exception, Q_t, is restricted to be a function of date $t-1$ and earlier information only. This informational constraint on Q_t reflects the limited participation feature of the model.

Both the informational restriction on the choice of Q_t and the cost of adjusting Q_t can be interpreted as ways of capturing, in a representative-agent model, more fundamental microeconomic frictions affecting household portfolio adjustment. For example, Caballero (1993), Marshall and Parekh (1994), and Schroder (1995) show that extremely small fixed costs of adjusting an economic choice variable at the individual level can imply extremely sluggish behavior of the corresponding macroeconomic aggregate. Even small costs of portfolio adjustment can imply a sluggish response of the aggregate household portfolio to monetary-policy shocks. Unfortunately, it is extremely difficult to formulate dynamic equilibrium models with cross-sectional heterogeneity that explicitly incorporate fixed adjustment costs. (The models of Caballero (1993), Marshall and Parekh (1994), and Schroder (1995) are all partial-equilibrium models.) Our formulation is an attempt to incorporate these effects into an equilibrium model in a tractable fashion.

4.3 *Firms*

We adopt a monopolistic competition framework, along the lines of Blanchard and Kiyotaki (1987). At time t, a final consumption good, Y_t is produced by a perfectly competitive firm. It does so by combining a continuum of intermediate goods, indexed by $i \in (0,1)$, using the technology:

$$Y_t = [\int_0^1 Y_{it}^{\frac{1}{\mu}} di]^{\mu}, \tag{17}$$

where $1 \leq \mu < \infty$ and Y_{it} denotes the time t input of intermediate good i.[21] Let P_t and P_{it} denote the time t price of the consumption good and intermediate good i, respectively. Profit maximization implies the Euler equation:

$$(\frac{P_t}{P_{it}})^{\frac{\mu}{\mu-1}} = \frac{Y_{it}}{Y_t}. \tag{18}$$

[21]The model with competitive and identical firms is a special case of this model in which μ is set to unity. This monopolistic competition paradigm is typically adopted in limited participation models because it implies a larger output response and a more realistic investment response to monetary-policy shocks. However, the responses of bond yields change very little when μ is set to unity.

Integrating (18) and imposing (17), we obtain the following relationship between the price of the final good and the price of the intermediate goods:

$$P_t = [\int_0^1 P_{it}^{\frac{1}{1-\mu}} di]^{(1-\mu)}. \tag{19}$$

Intermediate good i is produced by a monopolist who uses the following technology:

$$Y_{it} = \begin{cases} K_{it}^{\alpha} L_{it}^{1-\alpha} - \phi & \text{if } K_{it}^{\alpha} L_{it}^{1-\alpha} \geq \phi \\ 0 & \text{otherwise} \end{cases} \tag{20}$$

where $0 < \alpha < 1$. Here, L_{it} and K_{it} denote time t labor and capital used to produce the i^{th} intermediate good. The parameter ϕ denotes a fixed cost of production. We rule out entry and exit into the production of intermediate good i. Intermediate firms rent capital and labor in perfectly competitive factor markets. Economic profits are distributed to the firm's owner, the representative household.

The firm's choices are affected by monetary policy through a cash-in-advance constraint. Firms retain no cash from period to period, but workers must be paid in advance of production. As a result, firms need to borrow their wage bill, $W_t L_{it}$, from the financial intermediary at the beginning of the period. Repayment occurs at the end of time period t, at the gross interest rate, R_t. Given that the firm's only source of finance is through the financial intermediary, this feature of the model is one possible articulation of the credit channel in the monetary transmission mechanism.

Profit maximization leads the intermediate-good firm to set its price equal to a constant markup over marginal cost:

$$P_{it} = \mu MC_t, \tag{21}$$

implying

$$\frac{W_t R_t}{P_t} = \frac{f_{L,t}}{\mu}, \quad \frac{r_t}{P_t} = \frac{f_{K,t}}{\mu} \tag{22}$$

where $f_{L,t} = (1-\alpha)(K_{it}/L_{it})^{\alpha}$ is the marginal product of labor and $f_{K,t} = \alpha(L_{it}/K_{it})^{(1-\alpha)}$ is the marginal product of capital; and we have imposed the equilibrium condition, $P_{it} = P_t$ for all i.[22] Note that, in equation (22), the nominal interest rate R_t is determined in part by the marginal product of *labor*. This reflects the cash-in-advance constraint on labor inputs, described above. In equilibrium, all intermediate-goods firms choose the same labor and capital combinations, so we henceforth drop the i subscript.

[22] In deriving equation (22), we also use the following characterization of the marginal cost of the intermediate-good firm:

$$MC(r_t, R_t W_t) = \frac{1}{1-\alpha}(\frac{L_t}{K_t}^{\alpha})W_t R_t = \frac{1}{\alpha}(\frac{K_t}{L_t})^{(1-\alpha)} r_t.$$

4.4 *Financial intermediary*

At time t, a perfectly competitive financial intermediary buys and sells bonds with the household at prices $b_t^j, j = 1, ..., n$. The net amount of funds transferred from households to the financial intermediary is $M_t - Q_t$. The intermediary also receives a lump-sum cash injection, X_t, from the monetary authority. These funds are supplied to the loan market at the gross interest rate R_t. Demand in the loan market comes from the intermediate-good producers, who seek to finance their wage bill, $W_t L_t$. Clearing in the loan market requires:

$$W_t L_t = M_t - Q_t + X_t. \tag{23}$$

At the end of the period the intermediary pays off all maturing bonds B_t^1 to households, and distributes its profits (revenue from loan repayments minus the cost of paying off maturing bonds) to households as a dividend F_t:

$$F_t = R_t W_t L_t - B_t^1.$$

The assumption of perfect competition ensures that the nominal return earned by the household on one-period bonds equals R_t, the nominal return earned by the intermediary from its one-period loans to the firms. That is,

$$R_t = \frac{1}{b_t^1} \tag{24}$$

4.5 *Equilibrium*

Let λ_t, ν_t, and ξ_t denote the Lagrange multipliers associated with constraints (11), (12), and (13), respectively. The first-order conditions of the household are:

For Q_t :

$$E_{t-1}\{U_{H,t}H_t' \frac{1}{Q_{t-1}} - \beta U_{H,t+1}H_{t+1}' \frac{Q_{t+1}}{Q_t^2} + \nu_t + \xi_t + \lambda_t\} = 0. \tag{25}$$

For L_t:

$$U_{L,t} + (\nu_t + \xi_t)W_t = 0. \tag{26}$$

For C_t:

$$U_{C,t} = (\nu_t + \xi_t)P_t. \tag{27}$$

For K_{t+1}:

$$(\nu_t + \xi_t)P_t(K_{t+1} - (1-\delta)K_t) = \beta E_t\{\xi_{t+1}r_{t+1} + (\nu_{t+1} + \xi_{t+1})P_{t+1}(1-\delta)\} \tag{28}$$

For M_{t+1}:

$$\xi_t = \beta E_t\{\lambda_{t+1}\}. \tag{29}$$

For B_t^1:

$$\lambda_t b_t^1 = \xi_t. \tag{30}$$

For $B_t^j, j = 2, ..., n$:

$$\lambda_t b_t^j = \beta E_t\{\lambda_{t+1} b_{t+1}^{j-1}\}. \tag{31}$$

Notice that the conditional expectation in equation (25) is with respect to period $t-1$ information. This reflects the limited participation feature of the model. Using equations (27) and (29) we can eliminate multipliers ν_t and ξ_t. We can then use equations (22) and (24) to obtain the following equilibrium conditions:

$$E_{t-1}\{\lambda_t\} = E_{t-1}\{U_{H,t} H_t' \frac{1}{Q_{t-1}} - \beta U_{H,t+1} H_{t+1}' \frac{Q_{t+1}}{Q_t^2} + \frac{U_{C,t}}{P_t}\}. \tag{32}$$

$$U_{c,t} = \beta E_t\{\beta \lambda_{t+2} \frac{f_{K,t+1}}{\mu} P_{t+1} + U_{c,t+1}(1-\delta)\}. \tag{33}$$

$$U_{L,t} + U_{C,t} \frac{W_t}{P_t} = 0, \tag{34}$$

$$\frac{W_t R_t}{P_t} = \frac{f_{L,t}}{\mu}, \qquad \frac{r_t}{P_t} = \frac{f_{K,t}}{\mu}. \tag{35}$$

$$\lambda_t b_t^j = \beta^j E_t\{\lambda_{t+j}\}, \qquad j = 1, ..., n \tag{36}$$

4.6 *The term structure of nominal interest rates*

Equation (36) implies that the marginal utility of nominal wealth is λ_t, and the intertemporal marginal rate of substitution (IMRS) in nominal wealth is $\frac{\beta \lambda_{t+1}}{\lambda_t}$. This IMRS serves as the stochastic discount factor (in the sense of Hansen and Jagannathan (1991)) that determines the nominal bond yield of maturity j, denoted R_t^j, according to

$$R_t^j \equiv -\frac{1}{j} log[b_t^j] = -\frac{1}{j} log(E_t[\frac{\beta^j \lambda_{t+j}}{\lambda_t}]) \tag{37}$$

According to equation (37), the behavior of the term structure is completely determined by the stochastic process for λ_t. Furthermore, the yields of maturity $j > 1$ do not affect the determination of equilibrium λ_t or the other equilibrium prices or quantities. In other words, the equilibrium conditions are block-recursive between long bond yields and all other variables in the economy. The economic interpretation of this recursivity property is that longer-term bonds are *redundant assets*: If all bonds of maturity greater than one period were removed from the economy, equilibrium quantities and prices would be unchanged. As is typical in the asset-pricing literature, one can price redundant assets by first computing the process for the stochastic discount factor, and then use this process to compute asset prices.[23]

We implement this procedure as follows: Use equations (24), (32), (34), (35), and (36) evaluated at $j = 1$ to obtain the following expression for λ_t as a function of the processes for quantity and price variables:

$$\lambda_t = \frac{\beta(1-\alpha)}{\mu}\left(\frac{K_t}{L_t}\right)^{\alpha}\left(\frac{-U_{c,t}}{U_{L,at}}\right)\times$$

$$E_t\{U_{H,t+1}H'_{t+1}\frac{1}{Q_t} - \beta U_{H,t+2}H'_{t+2}\frac{Q_{t+2}}{Q^2_{t+1}} + \frac{U_{C,t+1}}{P_{t+1}}\} \tag{38}$$

In the absence of adjustment costs on Q_t (that is, if $H' = 0$), equation (32) implies that the conditional expectation of λ_t is simply the expected marginal utility of nominal consumption, as in the standard model. The marginal product of labor enters equation (38) due to the cash-in-advance constraint faced by the firms on labor inputs. To compute the bond yields, we first solve for the laws of motion of aggregate quantities, using techniques standard in the equilibrium business-cycle literature, and then use equations (37) and (38) to compute R^j_t. Details of the solution procedure we use can be found in the technical appendix.

4.7 *Calibration*

In choosing parameters for the model, we adhere closely to the equilibrium business-cycle literature. First, we choose $\alpha, \beta, \gamma, \psi, \mu, \delta$, and x as follows:

$$\alpha = 0.36, \quad \beta = 1.03^{-(1/12)}, \quad \gamma = 1,$$

$$\psi = 2/3, \quad \mu = 1.40, \quad \delta = 0.00667, \quad x = 0.00667.$$

The values of α (capital's share in the Cobb-Douglas production technology), β (the monthly subjective discount factor), and δ (the monthly capital depreciation rate) are standard choices. The value of γ implies a logarithmic specification in (16). The value of x implies a yearly monetary growth rate of 8%. The wage elasticity of labor supply in this model is $1/\psi$, so our choice of ψ implies a labor-supply elasticity of 1.5. This value is somewhat higher than most microeconomic estimates. For example, Card (1991), Killingsworth (1983), and Pencavel (1986) estimate elasticities near zero for males; whereas it is in the range of 0.5 to 1.5 for females, Killingsworth and Heckman (1986). However, the implied labor-supply elasticity in most

[23]Perhaps the most celebrated use of asset redundancy in a pricing model is the Black-Scholes formula. Other examples include Bansal and Viswanathan's (1993) nonlinear factor-pricing model and Duffie and Kan's (1996) affine model of the yield curve.

real business-cycle models substantially exceeds the value we use (for example, Christiano and Eichenbaum's (1992) model parameter estimates imply a Frisch labor-supply elasticity in excess of 5.0). The markup parameter μ is at the high end of the range used in the literature. Rotemberg and Woodford (1995) survey the evidence on markups and select a markup of this size.

We set ψ_0 to imply that in nonstochastic steady state, employment is unity. (This is a normalization that is without loss of generality.) In particular, we set

$$\psi_0 = \frac{\beta(1-\alpha)}{\mu(1+x)}\{\frac{\alpha\beta}{\mu(1+x)}\frac{1}{\frac{1}{\beta}+\delta-1}\}^{\frac{\alpha}{1-\alpha}}. \tag{39}$$

Our calibration implies that $\psi_0 = 2.95$. We set the fixed cost ϕ to imply that pure profits are zero in nonstochastic steady state, as follows:

$$\phi = (\frac{\mu-1}{\mu})K^{\alpha}, \tag{40}$$

where K denotes the nonstochastic steady-state stock of capital. Our calibration implies that $K = 181.45$, so $\phi = 1.86$.[24]

There is no literature to draw on in choosing values for the adjustment cost parameters c and d. Furthermore, the adjustment cost function $H(\cdot)$ is constructed so that both H and H' are zero in the nonstochastic steady state, so steady-state properties cannot be used to calibrate c and d. In our baseline calibration we choose c and d to imply a reasonable response of the one-month interest rate to a monetary-policy shock. In particular, we set $c = 2$ and $d = 1$. Finally, we choose the coefficients $\{\theta_i\}_{i=0}^{4}$ in the money-growth rule (10) to match the response of the monetary aggregate to a policy shock (as described in Section 4.1). The three sets of values are given in Table 2.

4.8 *Implications of the model*

Macroeconomic variables. Figure 5 displays responses of various macroeconomic aggregates other than the bond yields to a one-standard-deviation monetary contraction. These responses are qualitatively similar to the empirical responses displayed in Figure 1, and often correspond quantitatively as well.[25] The money-growth process is displayed in row four of Figure 5 and by construction is identical to the first five coefficients of the impulse

[24]Equation (40) implies that the ratio of the fixed cost ϕ to the steady-state output level is $\mu - 1$. Our parameterization therefore implies that, on average, 40% of output goes to pay the fixed cost.

[25]Output, employment, price, and money growth have been scaled up by a factor of 100 so the units are percent deviations from steady state. This makes the responses directly comparable to Figure 1.

Table 2:
Parameter Calibrations for the Money-Growth Process

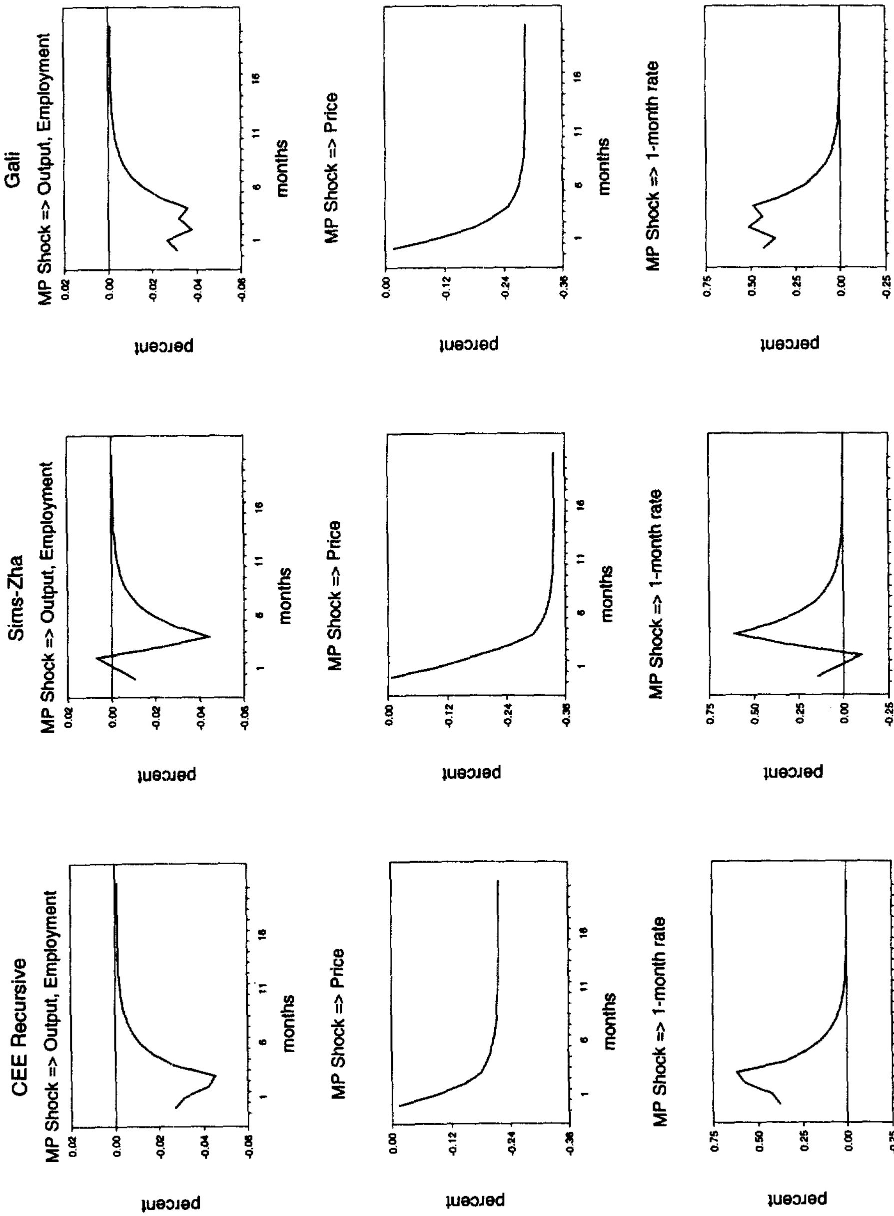
Identification Strategy	θ_0	θ_1	θ_2	θ_3	θ_4
CEE	.00041	.00081	.00067	.00038	-.00008
Gali	.00047	.00087	.00082	.00036	.00035
SZ	.00016	.00092	.00064	.00099	.00069

Notes for Table 2: This table gives the values of the moving average coefficients $\{\theta_i\}_{i=0}^{4}$ in the money-growth process (10) implied by our estimates of the Christiano, Eichenbaum, and Evans recursive strategy for identifying monetary-policy shocks (described in Section 2.2); the Sims-Zha nonrecursive identification strategy (described in Section 2.3); and the Gali identification strategy that incorporates long-run restrictions (described in Section 2.4).

response functions in row four of Figure 1. Depending on the calibration of the MA process, the price level declines to a level between 20% and 30% below the steady state. By comparison, the price responses in Figure 1 range between 15% (in the Christiano, Eichenbaum, and Evans identification) to 50% (in the Gali identification). However, these empirical responses do not appear to have attained the new steady state after 24 months, while the price response in the model is close to the new steady state after eight months.

As in the empirical results, the output level in the model declines in response to a contractionary monetary-policy shock. However, the response in the model is both smaller in magnitude and less persistent than in the data. In particular, the maximal output response in the model variants is a decline of about 0.045%, as compared to a decline of around 0.2% in the empirical exercises. Furthermore, the maximal response in the model is after three or four months, with the effect of the shock largely dissipated after ten months. In contrast, the point estimates in all three empirical exercises show a sustained response even after 24 months. As noted by Christiano, Eichenbaum, and Evans (1997a), the magnitude of the output response in a limited participation model of this type is determined largely by the elasticity of labor supply. (A monetary injection increases output by relaxing the firm's cash-in-advance constraint on labor inputs, and clearly the response of labor input to this increase in wages is critical.) For example, when we increase this elasticity to 2.5 (from our baseline calibration of 1.5), the maximal output response increases to 0.085%.

Finally, to see whether our specification of the adjustment cost function is reasonable, the last row of Figure 5 plots the response of the time-cost H_t of adjusting Q_t. (Since the steady-state labor supply is normalized to unity, the



CEE Recursive
Sims-Zha
Gali
MP Shock => Output, Employment
MP Shock => Price
MP Shock => 1-month rate
percent
months
0.02
0.00
-0.02
-0.04
-0.06
-0.12
-0.24
-0.36
0.75
0.50
0.25
-0.25
1
6
11
16

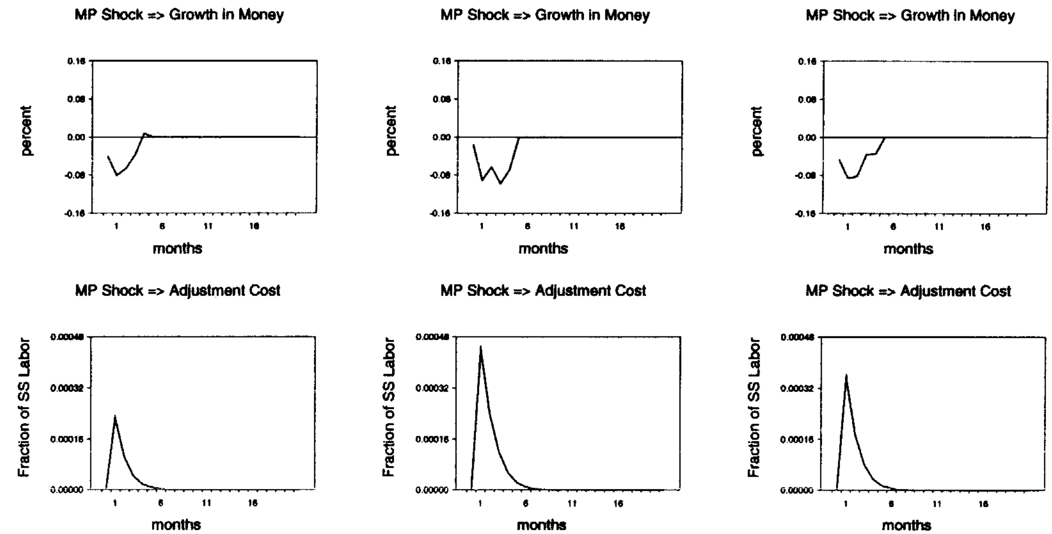
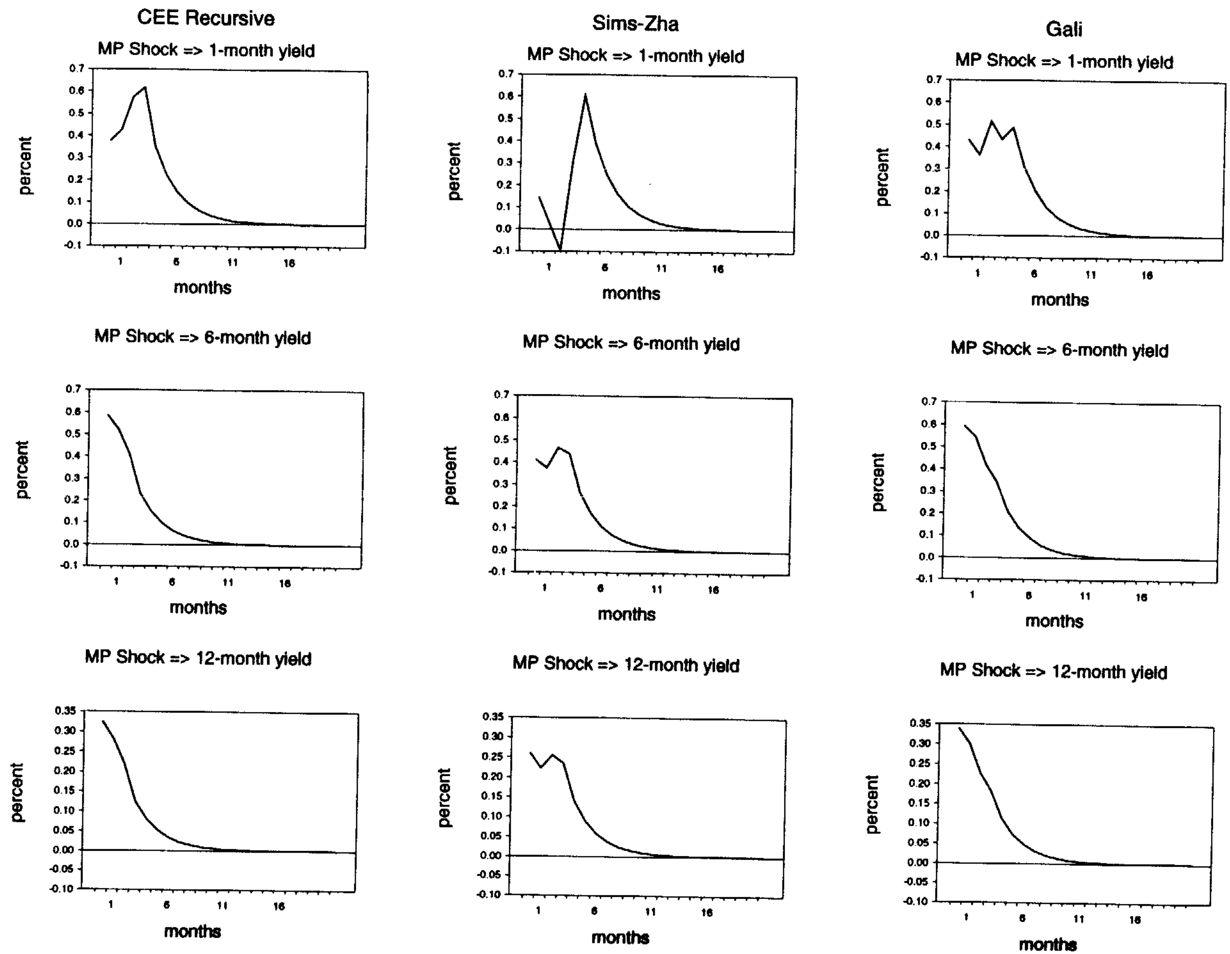


Figure 5

Responses of Macroeconomic Variables to Monetary-Policy Shocks in the Calibrated Limited Participation Model

Notes to Figure 5: This figure displays the responses to a one standard deviation contractionary monetary-policy shock, implied by the model described in Sections 4.1 - 4.4, of output, the price level, the one-month interest rate, the growth rate of money, and the adjustment cost of Q_t (as given in equation (14)). For output, the price level, and the growth rate of money, the responses are in units of percentage deviations from the steady state. For the one-month interest rate, the units are percentage points per annum. For the adjustment cost, the units are fractions of time devoted to employment in the steady state. The three columns calibrate the law of motion for money growth (equation (10)) to estimates from the CEE identification, the Sims-Zha identification, and the Gali identification, as given in Table 2. The remaining parameter calibrations are: $\alpha = 0.36$, $\beta = 1.03^{-(1/12)}$, $\sigma = 1$, $\psi = 2/3$, $\mu = 1.40$, $\delta = 0.00667$, $x = 0.00667$, $c = 2$, $d = 1$.


CEE Recursive
Sims-Zha
Gali
MP Shock => 1-month yield
MP Shock => 6-month yield
MP Shock => 12-month yield
percent
months

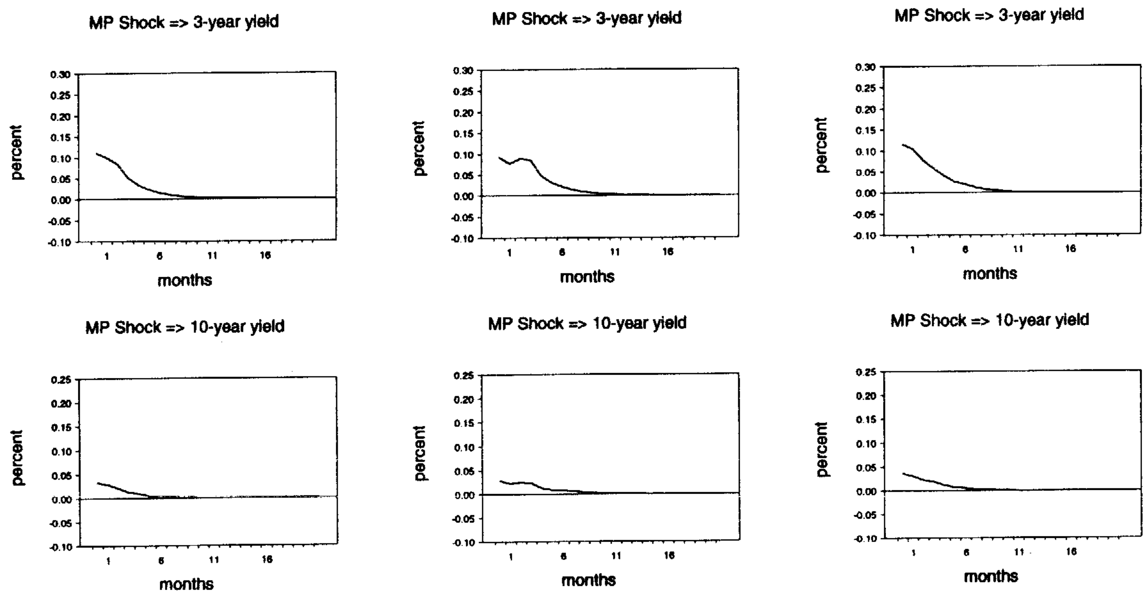


Figure 6

Responses of Bond Yields to Monetary-Policy Shocks in the Calibrated Limited Participation Model

Notes to Figure 6: The interpretation of Figure 6 is identical to that of Figure 5, except that the responses, implied by the model of Sections 4.1 - 4.4, of the continuously-compounded yields for zero-coupon bonds of maturities one month, six months, one year, three years, and ten years are displayed. The units are percentage points per annum. The calibration is as in Figure 5.

units for H_t are the fraction of steady-state labor used in adjusting Q_t.) As can be seen, the maximal adjustment cost engendered by the monetary-policy shock is in the period following the shock, and the magnitude is between 0.0002 and 0.0004 (depending on the model variant). We conclude that the costs implied by our specification are trivially small.

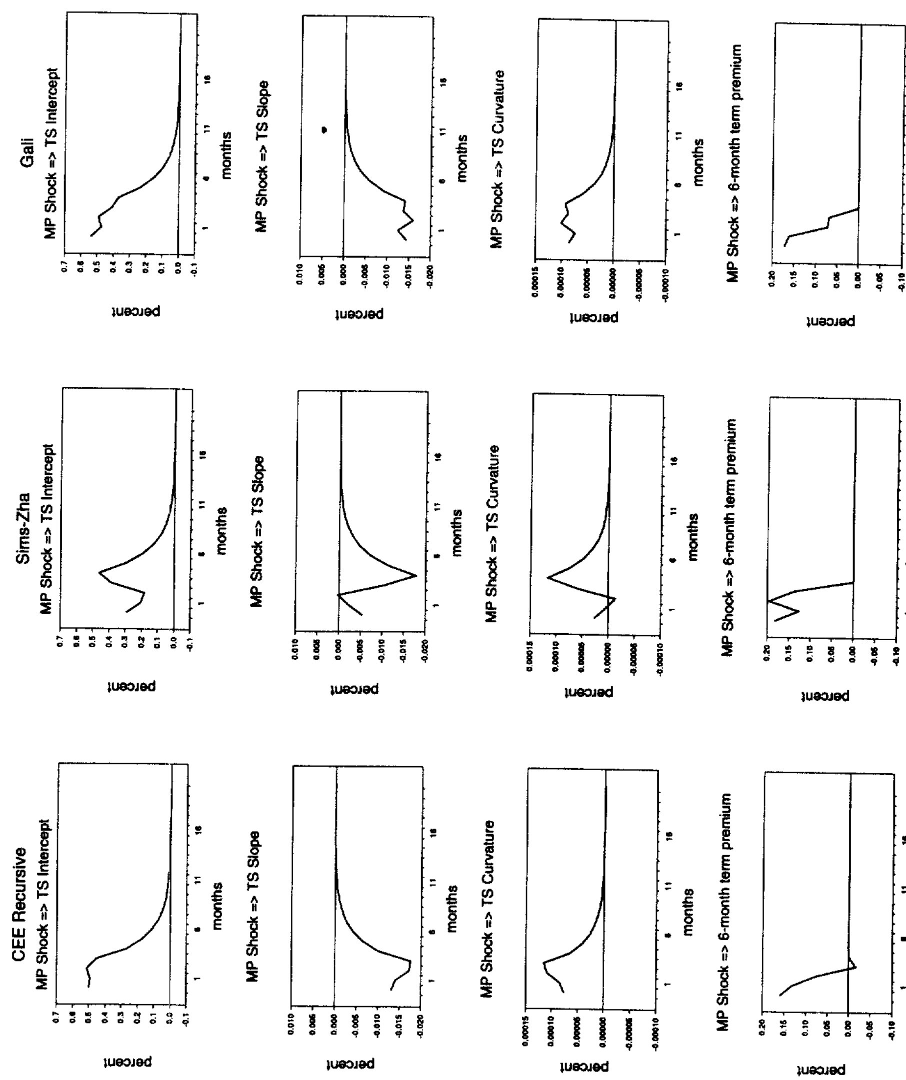
Bond yields. We find that the model replicates qualitatively the responses of bond yields to monetary-policy shocks. To preview our results, we find that a contractionary monetary-policy shock causes a positive but transient response in nominal bond yields, and an even bigger positive response in real bond yields. These responses are largest for the short-term yields, with the magnitude of the response declining as maturity increases. There is a positive response in the term premiums for the shorter-maturity bonds. In the rest of this section we describe these findings in greater detail.

In Figure 6 we display the responses of yields of different maturities to the contractionary monetary-policy shock. In presenting our model's results, the bond yield responses are annualized percentage point deviations from steady state. These responses are qualitatively similar to those found in the data, but the magnitudes are somewhat too large, at least for the shorter maturities. Consider the Christiano, Eichenbaum, and Evans calibration of the money-growth process: The contemporaneous response of the one-month yield in the model is 38 basis points, rising to 62 basis points three months after the shock. In contrast, our empirical exercises give a contemporaneous response of approximately 20 basis points, rising to a maximal response of 28 or 30 basis points one month later. The contemporaneous response in the model of the six-month yield is 59 basis points, which actually exceeds the contemporaneous response of the one-month yield. This is largely due to the persistence of the short-rate response in the model. (Recall that the expectations hypothesis would imply that the contemporaneous response of the six-month rate should equal the average of the first six-months' responses of the one-month rate.) However, the high contemporaneous response of the six-month rate is also due, in part, to the response of the six-month term premium, as discussed below. The remaining yields' responses decline in magnitude as maturity increases, much as in the empirical exercises. The initial responses of the three- and ten-year yields are 11 basis points and 3 basis points, respectively. These numbers are rather close to the responses found in the data. The responses implied by the Gali calibration are similar to those of the Christiano, Eichenbaum, and Evans calibration. The responses of the one-month yield implied by the Sims-Zha calibration differ somewhat from the other two, in that the contemporaneous response to the policy shock is rather small (about 15 basis points). Still, the maximal response in this calibration (which occurs four months after the shock) is over 60 basis points, substantially exceeding any of the point estimates in the empirical exercises.

Shape of the term structure. Figure 7 displays the model's implications for the *intercept, slope*, and *curvature* term-structure descriptors. These responses can be computed directly from the individual bond-yield responses as follows. Consider a bond yield of maturity j months, and let R_i^j denote its response i months following a monetary-policy shock. The i^{th} response of the three descriptors—*intercept, slope*, and *curvature*—can be computed by projecting R_i^j on a constant, j, and j^2. For each of the model parameterizations reported, the *intercept* increases for about four months and falls rapidly. The maximal *intercept* responses are about 50 basis points, whereas in the data (Figure 3), the maximal responses are around 18 basis points. This result highlights again that the model's implications for short-maturity bond yields exceed the empirical responses. The *slope* response falls, consistent with the declining influence of the monetary-policy contraction on longer-maturity yields. The quantitative responses of the model and data are quite similar; for example, the Christiano, Eichenbaum, and Evans impact response in Figure 3 is -0.015 and the model's impact response is -0.013. Finally, the model's implication for *curvature* is qualitatively consistent with the empirical responses in Figure 3. However, the magnitude of this effect in the model is severely diminished relative to the data.

Term premiums. The empirical exercises provide some evidence that term premiums increase following a contractionary monetary-policy shock. Our model does imply a positive term-premium response. As shown in Figure 7, the six-month term premiums in all three variants of the model display a positive response of 16-18 basis points. This is somewhat larger than the point estimates in our empirical exercises (in which the maximal response was 8 - 10 basis points). Furthermore, the term-premium response in the model dissipates rapidly, becoming essentially zero after three or four months. In contrast, the response of the six-month term premium in the empirical exercises appears to increase over the first four or five months. The response of term premiums declines monotonically with maturity, being 3 basis points for the three-year premium and less than 1 basis point for the ten-year premium.

Real yields. As in the data, we find that the real yields display a more pronounced response than nominal yields. (This follows immediately from the negative response of the inflation rate.) Figure 8 displays the real yield responses. The responses of shorter-term real yields in the model are larger and less persistent than in the data. For example, the maximal response of the six- and twelve-month real yields in the Christiano, Eichenbaum, and Evans calibration of the model are 97 basis points and 53 basis points, respectively. In the data, the maximal responses from the Christiano, Eichenbaum, and Evans and Sims and Zha monetary-policy shocks are under 40 basis points. The Gali identification produces a maximal response of 45 points (recall that the Gali monetary-policy shocks generate the largest fall in the price level).


CEE Recursive
Sims-Zha
Gali
MP Shock => TS Intercept
MP Shock => TS Slope
MP Shock => TS Curvature
MP Shock => 6-month term premium
percent
months

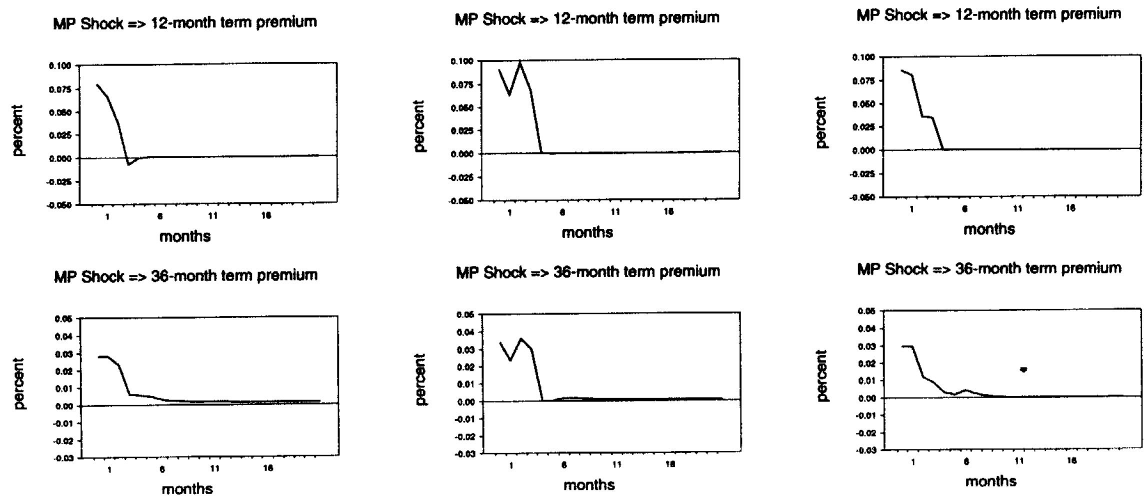
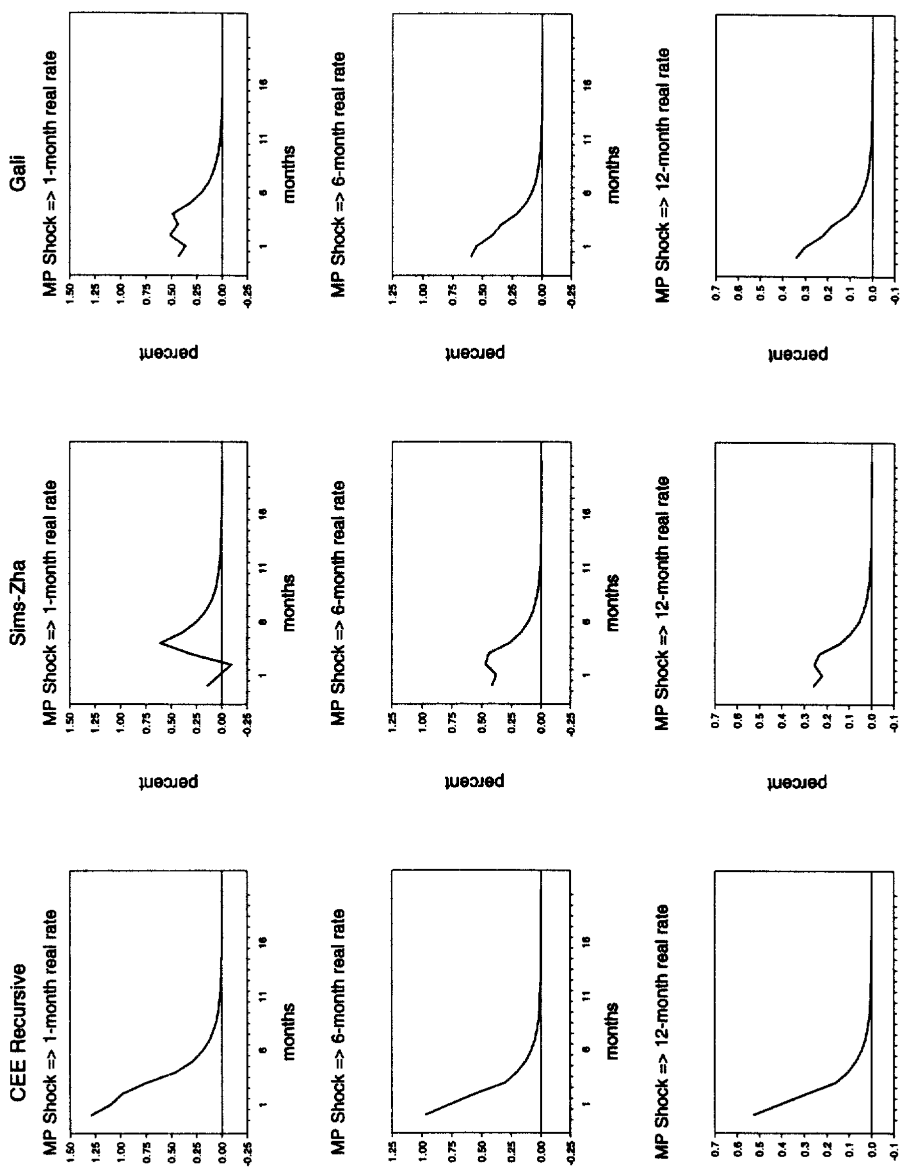


Figure 7

Responses of Term Structure Descriptors and Term Premiums to Monetary-Policy Shocks in the Calibrated Limited Participation Model

Notes to Figure 7: The interpretation of Figure 7 is identical to that of Figure 5, except that the responses, implied by the model of Sections 4.1 - 4.4, of the three term structure descriptors *intercept, slope,* and *curvature* (rows 1-3) and of the six-month, twelve-month, and 36-month term premiums are displayed. The three term structure descriptors are computed as described in Section 3.1. The *j*-period term premium, TP_t^j is defined in equation (9). The calibration is as in Figure 5.


CEE Recursive
Sims-Zha
Gali
MP Shock => 1-month real rate
MP Shock => 6-month real rate
MP Shock => 12-month real rate
percent
months

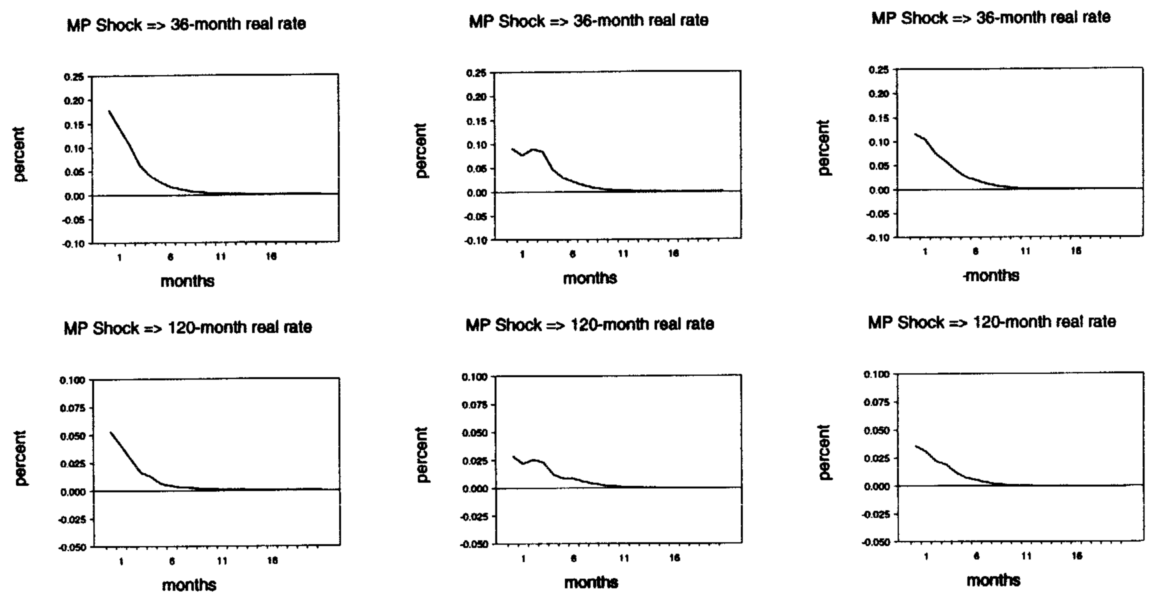
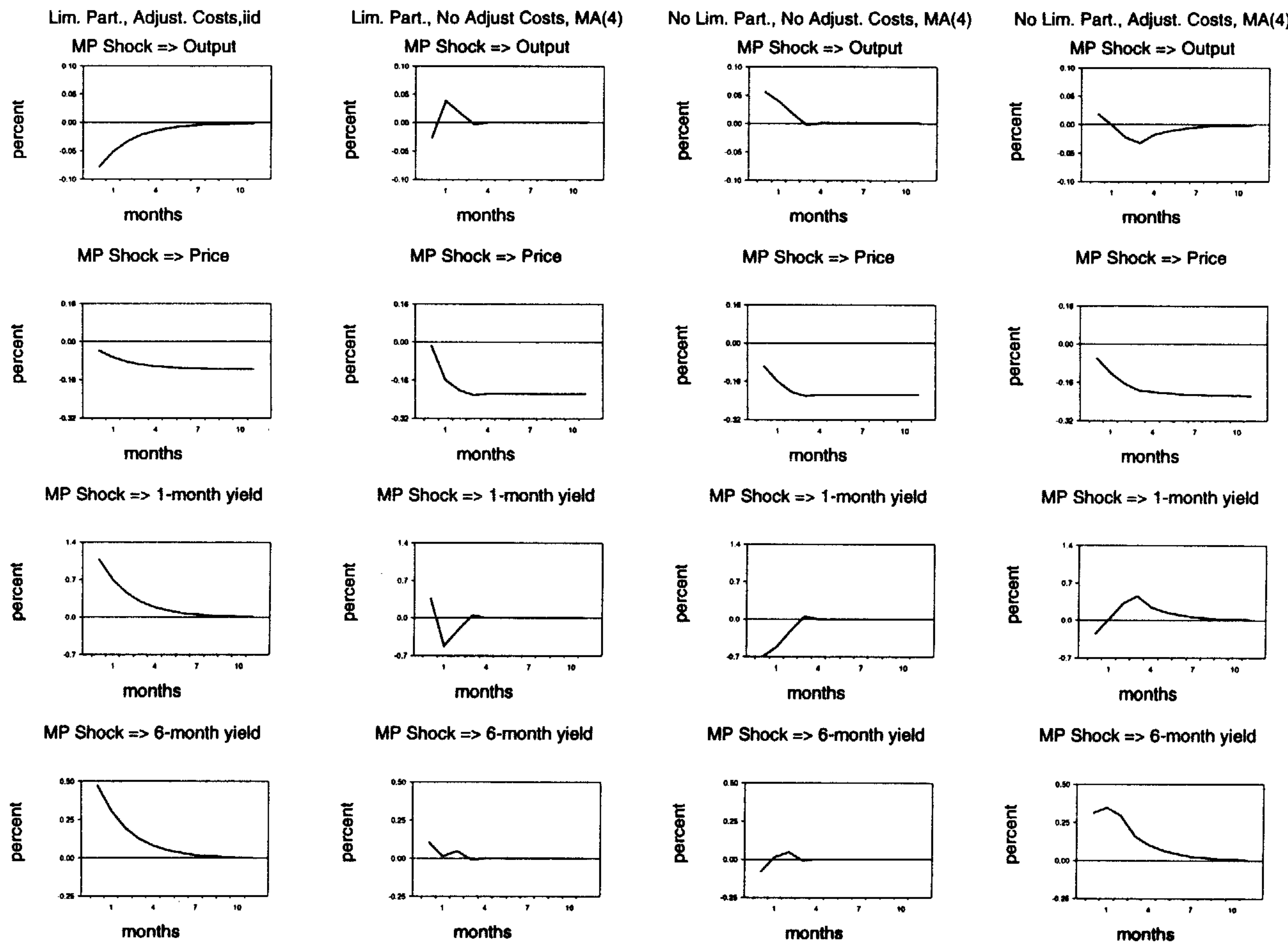


Figure 8

Responses of Real Interest Rates to Monetary-Policy Shocks in the Calibrated Limited Participation Model

Notes to Figure 8: The interpretation of Figure 8 is identical to that of Figure 5, except that the responses, implied by the model of Sections 4.1 - 4.4, of the real bond yields of maturities one month, six months, twelve months, three years, and ten years are displayed. The units are percentage points per annum. The calibration is as in Figure 5.


Lim. Part., Adjust. Costs,iid
Lim. Part., No Adjust Costs, MA(4)
No Lim. Part., No Adjust. Costs, MA(4)
No Lim. Part., Adjust. Costs, MA(4)
MP Shock => Output
MP Shock => Price
MP Shock => 1-month yield
MP Shock => 6-month yield
percent
months

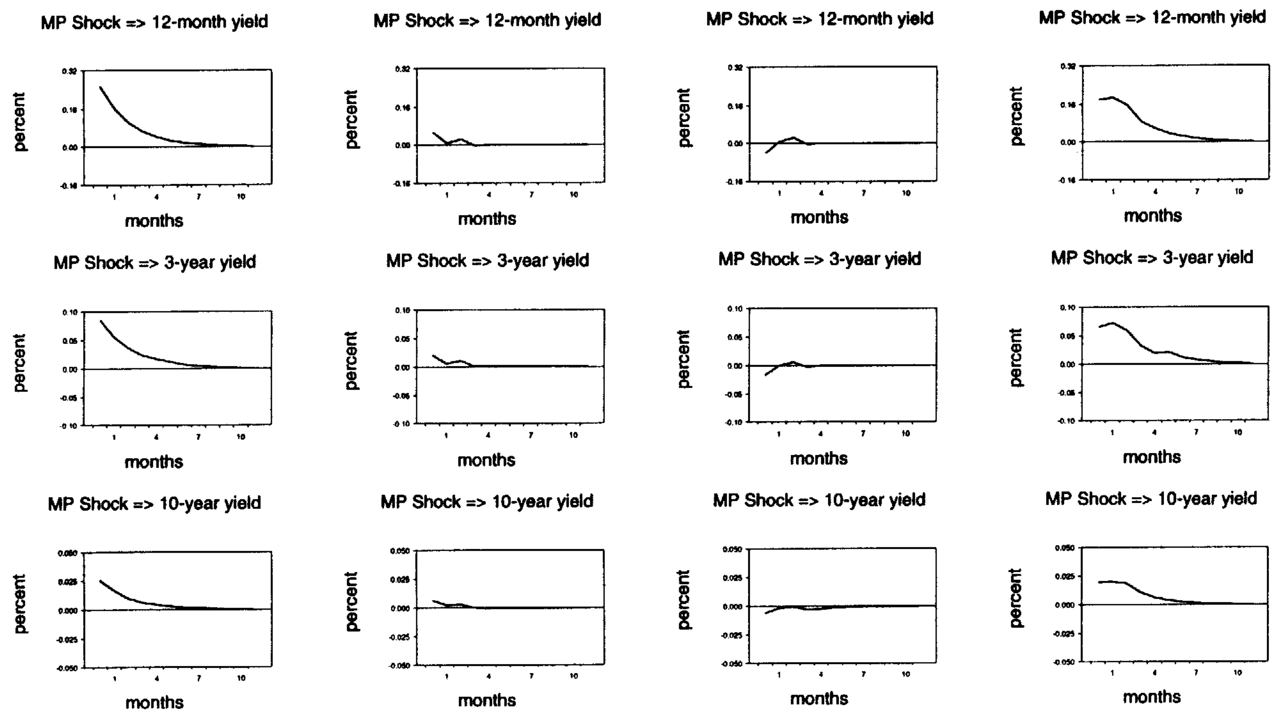
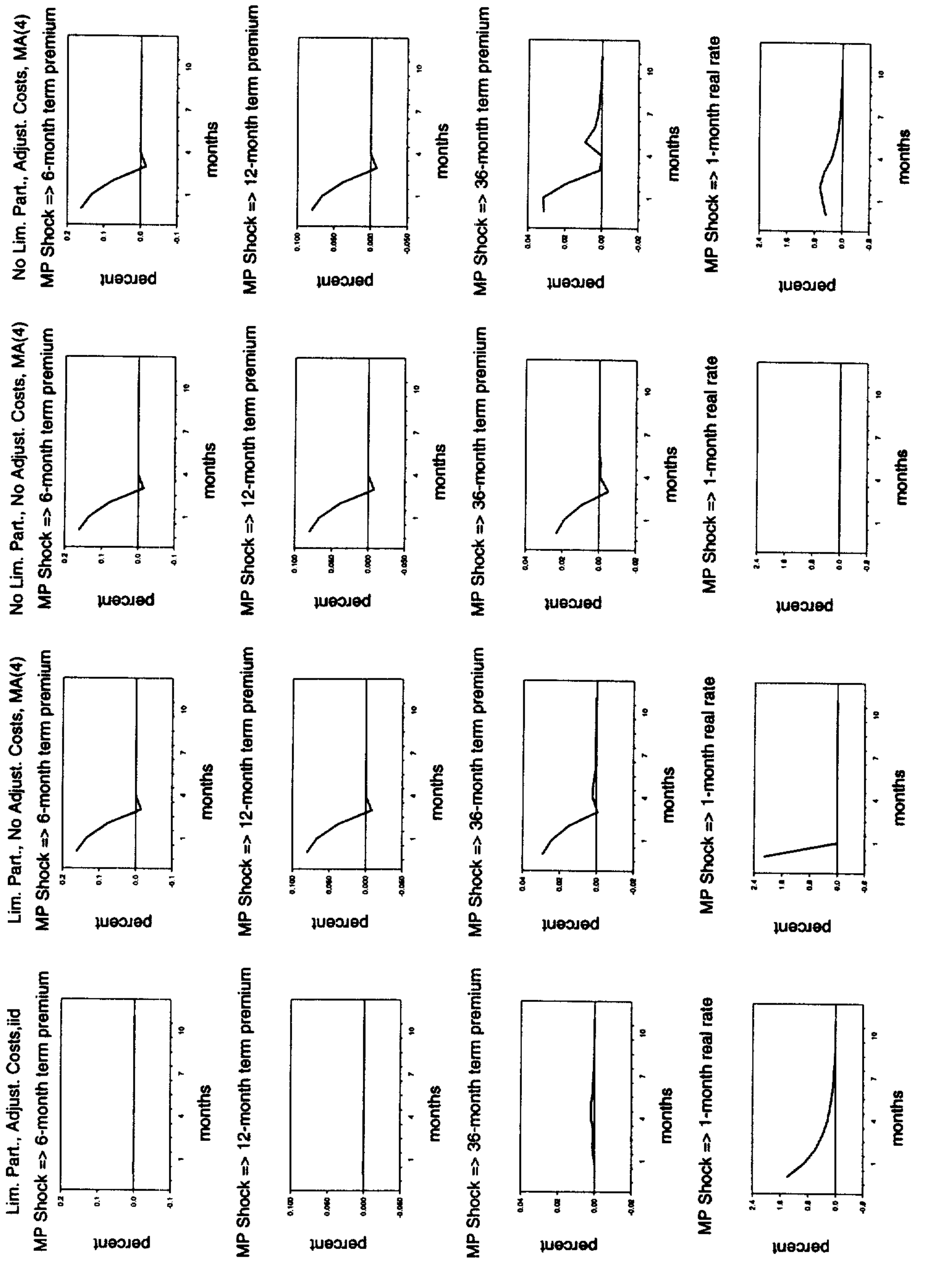


Figure 9: Part One


Lim. Part., Adjust. Costs,iid
Lim. Part., No Adjust. Costs, MA(4)
No Lim. Part., No Adjust. Costs, MA(4)
No Lim. Part., Adjust. Costs, MA(4)
MP Shock => 6-month term premium
MP Shock => 12-month term premium
MP Shock => 36-month term premium
MP Shock => 1-month real rate
percent
months

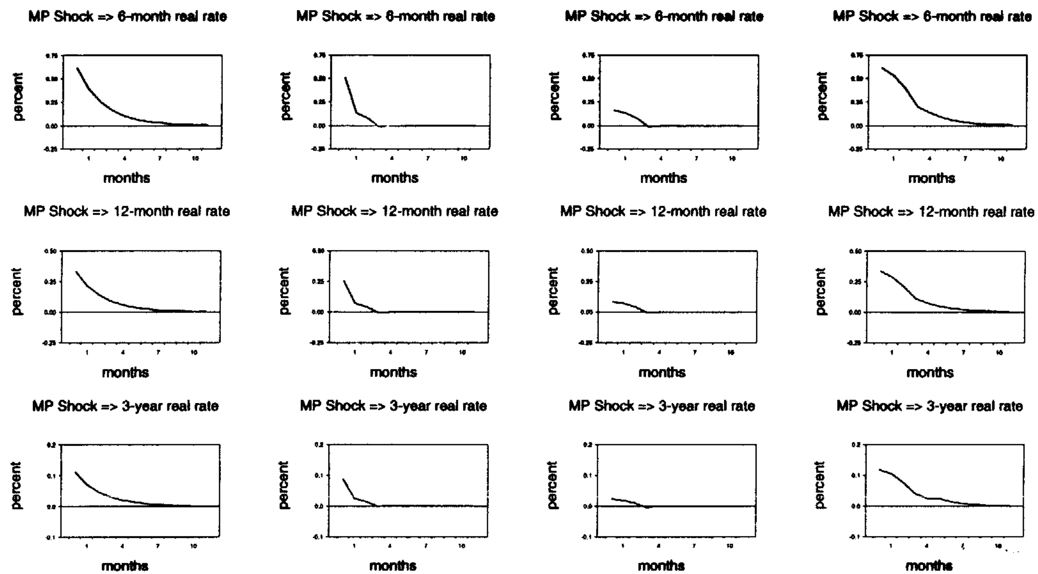


Figure 9: Part Two

Impulse Responses Implied by Variants of the Model of Section 4

Notes to Figure 9: The responses implied by four simple variants of the model of Sections 4.1 - 4.4 are displayed for the following variables: output (row one); the price level (row two); bond yields of maturities one month, six months, twelve months, three years, and ten years (rows three through seven); term premiums of maturities six months, twelve months, and three years, as defined in equation (9) (rows eight through ten); and real bond yields of maturities one month, six months, twelve months, and three years (rows 11 through 14). The simple model variants are as follows. Column 1: baseline model of Figure 5, but with money growth i.i.d.; Column 2: baseline model of Figure 5, but with adjustment cost set equal to zero; Column 3: baseline model of Figure 5, but without the limited participation constraint, and with adjustment cost set equal to zero; Column 4: baseline model of Figure 5 without the limited participation constraint, but with adjustment costs as in Figure 5.

The maximal response in the model is the contemporaneous response, while the maximal response in the data occurs four to five months after the date of the shock. The response of the three-year yield in the model (18 basis points) is close to that in the empirical exercises, so the excessive responses in the model appear to be confined to the shorter-term yields.

4.9 *Sources of the liquidity effect in the model*

It is of interest to see which elements of the model account for its ability to replicate the qualitative features found in our empirical exercises. To this end, we experimented with simpler versions of the model. These results are displayed in Figure 9. First, suppose money growth were i.i.d., rather than MA(4). That is, we replace equation (10) with

$$x_t = x + \tilde{\theta}_0 \epsilon_t \tag{41}$$

where, to ensure that x_t has the same variance as before, we set $\tilde{\theta}_0 = \sqrt{\Sigma_{i=0}^4 \theta_i^2}$. The results of this exercise are in the first column of Figure 9. Not surprisingly, the general response patterns are quite similar, although the dynamics are less complex. (Basically, all responses appear to die off exponentially.) The price-level response is smaller in the i.i.d. case, but this is purely an artifact of our normalizing the variances to be equal.[26] The most notable effect of removing serial correlation in money growth is to completely eliminate the term-premium responses. This in turn reduces the responses of the longer nominal yields, and (along with the reduced inflation response) reduces the responses of the real yields.

The second column of Figure 9 keeps the MA(4) structure in equation (10) (calibrated to the Christiano, Eichenbaum, and Evans identification), but sets the Q_t adjustment cost (in equation (14)) to zero. By construction, the impact responses of all variables to the contractionary monetary-policy shock are identical to those in the full model. However, the response patterns following the date of the shock differ substantially. In particular, the smooth responses in our baseline calibration are replaced by a rapid reversal of the contemporaneous response. Consider the output response. The initial liquidity effect reduces output, but thereafter the expected inflation effect dominates. That is, the contractionary shock signals that future inflation will be lower. (Indeed, our estimated MA process implies a greater reduction in money growth the period after the shock is observed.) Inflation acts like a distortionary tax, so reduced expected inflation tends to increase output.

[26]That is, $\sqrt{\Sigma_{i=0}^4 \theta_i^2} < \Sigma_{i=0}^4 \theta_i$, so while the unconditional variance of x_t is the same for both processes, the cumulative effect of a monetary shock on the money stock (and therefore the price level) is greater in the MA(4) specification.

We next explore the role of the limited participation feature in the model. Consider first a model *without* the limited participation features – that is, the household now observes the monetary-policy shock before choosing Q_t. Furthermore, we fix the adjustment cost a zero, and the money-growth rate is i.i.d., as in equation (41). This model now behaves as a conventional cash-in-advance model with i.i.d. shocks. The model's responses are completely neutral. No endogenous variables (other than the price level) respond to the policy shock. That is, there is neither a liquidity effect nor an expected inflation effect.

The third column in Figure 9 suppresses both the limited participation constraint and the adjustment costs, but assumes that the money-growth rate follows the MA(4) process estimated from the Christiano, Eichenbaum, and Evans identification. Now, the only effect of a monetary-policy shock is the expected inflation effect: As a result, the contractionary monetary shock causes nominal interest rates to decline. That is, the impact effect of a contractionary monetary-policy shock on bond yields has a counterfactual response relative to Figure 1. As described above, output responds positively to this reduced expected inflation. Interestingly, the term-premium responses are virtually identical to those found in the baseline calibration. Evidently, the term-premium responses are driven not by the limited participation feature of the model, but purely by the serial correlation in the money-growth process. As a result of this positive term-premium response, the response of the intermediate-term nominal yields turns positive after the initial decline.

In the fourth column of Figure 9 we include the Q_t-adjustment cost (as in equation (14)) to the model in the third column, but we continue to suppress the limited participation constraint. The adjustment cost not only smooths the impulse responses, but also attenuates the contemporaneous response: Households do not reduce their demand for Q_t-money as much as they would in the absence of adjustment costs, so less funds flow from the household to the financial intermediary. On impact the variables other than the long yields respond like the cash-in-advance model of column three. After the initial impact, the model's responses are like the baseline calibration, since unanticipated money movements are absent. The impact response of the j-period long yield incorporates the average of the first j responses of the one-month yield, which is positive for moderately large j. As a result, the impact response of the longer yields is positive. Of course, the positive term-premium response magnifies the impact on the long yields.

Perhaps the most curious result from these exercises is that the term-premium response arises neither from the adjustment costs nor from the limited participation constraint, but purely from the serial correlation in the money-growth process. More generally, it is the information content in the monetary-policy shock that causes term premiums to respond. Term premiums are a reflection of time-varying conditional higher moments in the logarithm of the marginal utility of wealth[27] (our variable λ_t). While the exogenous shocks in this model are conditionally homoskedastic, $\log(\lambda_t)$ responds nonlinearly to these shocks. These nonlinearities in turn can induce conditionally heteroskedastic behavior in $\log(\lambda_t)$. Evidently, a monetary-policy shock conveys not only information about the conditional first moment of future money growth, but also information about conditional higher moments of $\log(\lambda_t)$.

5 Conclusions

The results of this paper are straightforward and quite intuitive. We find clear evidence that the short-term effect of monetary policy takes the form of a liquidity effect: A monetary contraction raises interest rates while reducing expected inflation, inducing a pronounced rise in real interest rates. This effect is rather transitory, dissipating between six to twelve months after the initial impulse. Monetary-policy shocks primarily affect short-term interest rates, with a diminishing effect on longer-term rates. Much of the response of longer-term rates can be explained by the expectations hypothesis. There is some weak evidence that the response of the shorter-maturity bonds is enhanced by a positive response of term premiums. Finally, most of the variance of interest rates is due to sources other than monetary policy. In particular, monetary policy is a nontrivial source of the short-run variability of short-term rates, but it represents a negligible source of variability for long-term rates. We are encouraged by the apparent robustness of these results: These basic patterns emerge under three rather different identification approaches.

We also find that a simple dynamic equilibrium model, in which nominal rigidities take the form of a limited participation constraint, is consistent with the broad patterns we have detected in the data. This suggests that our empirical evidence is in no way anomalous, but has a simple equilibrium explanation. We believe that models of this sort may help us understand the interaction between monetary policy and asset markets generally. Having said this, we note that the model does require some degree of sluggishness in household portfolio adjustment. We have modeled this by imposing a simple adjustment cost. It would be more satisfactory, from a theoretical standpoint, to be explicit about the microeconomic frictions that underlie this slow aggregate portfolio adjustment.

We also note that our model only includes monetary-policy shocks. In principle, nonlinearities of the type we encountered in pricing the zero-coupon

[27]For a formal derivation of this result, see Bekaert, Hodrick, and Marshall (1997a).

bonds imply that the impulse responses to monetary shocks need not be invariant to the presence of other shocks in the model. It is an open question whether these nonlinearities substantially affect the economic analysis. It would be of interest to explore the role of monetary policy in a model of interest rates that incorporates a full set of exogenous impulses, such as technology shocks, preference shocks, and shocks to the transactions technology. (This last type of shock induces exogenous movements in money demand.) In addition, a model of this sort can be used to explore the effect of changes in the monetary-policy reaction function. Such an analysis would surely encounter a host of issues not discussed here.

6 Technical appendix: solving the model

We solve the model using the partial linearization method described in Christiano, Eichenbaum, and Evans (1997c). Let us define $q_t = Q_t/M_t$. The variable q_t is stationary. The model has two endogenous state variables: K_{t+1} and q_t; the exogenous state variables are $[x_t, \epsilon_t, \epsilon_{t-1}, \epsilon_{t-2}, \epsilon_{t-3}]$. We can reduce the equilibrium conditions of the model to two Euler equations involving only the processes for these state variables. A linear approximation to the integrands of these Euler equations is taken, and the resulting system of stochastic difference equations is solved in the usual manner to yield a linear law of motion for the endogenous state vector $[K_{t+1}, q_t]'$ (regarded now as differences from the steady-state values):

$$\begin{bmatrix} K_{t+1} \\ q_t \end{bmatrix} = A \begin{bmatrix} K_t \\ q_{t-1} \end{bmatrix} + B[x_t, \epsilon_t, \epsilon_{t-1}, \epsilon_{t-2}, \epsilon_{t-3}]' \tag{42}$$

where A and B are coefficient matrices of the appropriate dimensions. All other variables of interest are known functions of the processes for $[K_t, q_{t-1}]'$ and $[x_t, \epsilon_t, \epsilon_{t-1}, \epsilon_{t-2}, \epsilon_{t-3}]'$, so they can be computed exactly, once laws-of-motion (10) and (42) are known. That is, no other linear approximations are used other than equation (42).

In order to compute j-period bond yields, the conditional expectation $E_t[\lambda_{t+j}]$ must be evaluated. For j equal to twelve months or less we do so by Gauss-Hermite quadrature, using a two-point discretization[28] over $\{\epsilon_{t+1}, \ldots, \epsilon_{t+j}\}$. For j equal to 36 months or 120 months, the quadrature procedure is computationally infeasible, so we use a Monte Carlo method. In particular, we simulate a time series $\{\lambda_t, K_t, q_{t-1} x_t, \epsilon_t\}_{t=1}^{50,000}$, and we regress λ_{t+j} on a third-order Chebyshev polynomial function of $[K_t, q_{t-1}, x_t, \epsilon_t, \epsilon_{t-1}, \epsilon_{t-2}, \epsilon_{t-3}]'$, using the fitted regression as an approximation for $E_t[\lambda_{t+j}]$. As a check

[28]When the order of the discretization is increased to six, the implications for the impulse responses we study are virtually unaffected.

on the accuracy of this procedure, we compute the impulse responses for the twelve-month yield using both the Monte Carlo procedure and the quadrature procedure. The resulting impulse responses are virtually indistinguishable.

References

Backus, D., Gregory, A., and Zin, S., (1989). Risk Premiums in the Term Structure: Evidence from Artificial Economies. *Journal of Monetary Economics*, **24**: 371-99.

Bansal, R. and Viswanathan, S., (1993). No Arbitrage and Arbitrage Pricing: A New Approach. *Journal of Finance*, **48**: 1231-62.

Bekaert, G., Hodrick, R, and Marshall, D., (1997a). 'Peso Problem' Explanations for Term Structure Anomalies. Federal Reserve Bank of Chicago Working Paper, WP-97-7.

Bekaert, G., Hodrick, R., and Marshall, D., (1997b). The Implications of First-Order Risk Aversion for Asset Market Risk Premiums. *Journal of Monetary Economics*, **40**: 3-41.

Bernanke, B., (1986). Alternative Explanations of the Money-Income Correlation. *Carnegie-Rochester Conference Series on Public Policy*, **25**: 49-99.

Bernanke, B. and Blinder, A., (1992). The Federal Funds Rate and the Channels of Monetary Transmission. *American Economic Review*, **82**: 901-921.

Bernanke, B., Gertler, M., and Watson, M., (1997). Systematic Monetary Policy and the Effects of Oil Price Shocks. *Brookings Papers on Economic Activity*, **1**: 91-142.

Bernanke, B. and Woodford, M., (1997). Inflation Forecasts and Monetary Policy. *Journal of Money, Credit, and Banking*, **29**: 653-684.

Blanchard, O. and Kiyotaki, N., (1987). Monopolistic Competition and the Effects of Aggregate Demand. *American Economic Review*, **77**: 647-666.

Blanchard, O. and Quah, D., (1989). The Dynamic Effects of Aggregate Demand and Supply Disturbances. *American Economic Review*, **79**: 655-673.

Blanchard, O. and Watson, M., (1986). Are Business Cycles All Alike? *The American Business Cycle: Continuity and Change*, (ed.) R. Gordon. Chicago: University of Chicago Press.

Bliss, R., (1994). Testing Term Structure Estimation Methods. Manuscript, Federal Reserve Bank of Atlanta.

Caballero, R., (1993). Durable Goods: An Explanation for Their Slow Adjustment. *Journal of Political Economy*, **101**: 351-384.

Card, D., (1991). Intertemporal Labor Supply: An Assessment. Working paper, National Bureau of Economic Research No. 3602.

Christiano, L. and Eichenbaum, M., (1992). Current Real-Business-Cycle Theories and Aggregate Labor-Market Fluctuations. *American Economic Review*, **82**: 430-50.

Christiano, L. and Eichenbaum, M., (1995). Liquidity Effects, Monetary Policy and the Business Cycle. *Journal of Money, Credit and Banking*, **27**: 1113-1136.

Christiano, L., Eichenbaum, M., and Evans, C., (1996a). The Effects of Monetary-Policy Shocks: Evidence from the Flow of Funds. *Review of Economics and Statistics*, **78**: 16-34.

Christiano, L., Eichenbaum, M., and Evans, C., (1996b). Identification and the Effects of Monetary-Policy Shocks. *Financial Factors in Economic Stabilization and Growth*, (eds.) M. Blejer, Z. Eckstein, Z. Hercowitz, and L. Leiderman. Cambridge: Cambridge University Press.

Christiano, L., Eichenbaum, M., and Evans, C., (1997a). Sticky Price and Limited Participation Models: A Comparison. *European Economic Review*, **41**: 1201-1249.

Christiano, L., Eichenbaum, M., and Evans, C., (1997b). Monetary-Policy Shocks: What Have We Learned and to What End? Forthcoming in *Handbook of Macroeconomics*, (eds.) J. Taylor and M. Woodford. Amsterdam: North-Holland.

Christiano, L., Eichenbaum, M., and Evans, C., (1997c). Modeling Money. Working Paper, National Bureau of Economic Research No. 6371.

Cook, T. and Hahn, T., (1989). The Effect of Changes in the Federal Funds Target on Market Interest Rates in the 1970s. *Journal of Monetary Economics*, **24**: 331-351.

Dai, Q. and Singleton, K., (1997). Specification Analysis of Affine Term

Structure Models. Working Paper, National Bureau of Economic Research No. 6128.

den Haan, W., (1995). The Term Structure of Interest Rates in Real and Monetary Economics. *Journal of Economic Dynamics and Control*, **19**: 909-40.

Duffie, D. and Kan, R., (1996). A Yield-Factor Model of Interest Rates. *Mathematical Finance*, **6**: 379-406.

Eichenbaum, M., (1992). Comment on "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy." *European Economic Review*, **36**: 1001-1011.

Estrella, A. and Hardouvelis, G., (1991). The Term Structure as a Predictor of Real Economic Activity. *Journal of Finance*, **46**: 555-576.

Fama, E. and French, K., (1988). Permanent and Temporary Components of Stock Prices. *Journal of Political Economy*, **96**: 246-273.

Fuerst, T., (1992). Liquidity, Loanable Funds, and Real Activity. *Journal of Monetary Economics*, **29**: 3-24.

Gali, J., (1992). How Well Does the IS-LM Model Fit Post War Data? *Quarterly Journal of Economics*, **107**: 709-738.

Goodfriend, M. and King, R., (1997). The New Neoclassical Synthesis and the Role of Monetary Policy. *NBER Macroeconomics Annual*, 231-283.

Gordon, D. and Leeper, E., (1994). The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification. *Journal of Political Economy*, **102**: 1228-1247.

Hansen, L. and Jagannathan, R., (1991). Implications of Security Market Data for Models of Dynamic Economics. *Journal of Political Economy*, **99**: 225-62.

Hardouvelis, G., (1994). The Term Structure Spread and Future Changes in Long and Short Rates in the G7 countries: Is There a Puzzle? *Journal of Monetary Economics*, **33**: 255-284.

Jorion, P. and Mishkin, F., (1991). A Multicountry Comparison of Term-

Structure Forecasts at Long Horizons. *Journal of Financial Economics*, **29**: 59-80.

Killingsworth, M., (1983). *Labor Supply.* New York: Cambridge University Press.

Killingsworth, M. and Heckman, J., (1986). Female Labor Supply: A Survey. *Handbook of Labor Economics*, (eds.) O. Ashenfelter and R. Layard. Amsterdam: North-Holland.

Knez, P., Litterman, R., and Scheinkman, J., (1994). Explorations into Factors Explaining Money Market Returns. *Journal of Finance*, **49**: 1861-1882.

Leeper, E., Sims, C., and Zha, T., (1996). What Does Monetary Policy Do? *Brookings Papers on Economic Activity*, **2**: 1-63.

Litterman, R. and Scheinkman, J., (1991). Common Factors Affecting Bond Returns. *The Journal of Fixed Income*, **1**: 54-61.

Lucas, R., (1990). Liquidity and Interest Rates. *Journal of Economic Theory*, **50**: 237-64.

McCallum, B., (1994a). Monetary Policy and the Term Structure of Interest Rates. Working Paper, National Bureau of Economic Research No. 4938.

McCallum, B., (1994b). A Semi-Classical Model of Price-Level Adjustment. *Carnegie-Rochester Conference Series on Public Policy*, **41**: 251-284.

McCulloch, J., (1990). The Term Structure of Interest Rates: Appendix B: U.S. Term Structure Data, 1946-87. *Handbook of Monetary Economics*, 1, (eds.) B. Friedman and F.H. Hahn. Amsterdam: North-Holland.

McCulloch, J. and Kwon, H., (1993). U.S. Term Structure Data, 1947-199. Ohio State University Working Paper, 93-6.

Marshall, D. and Parekh, N., (1994). The Effect of Costly Consumption Adjustment on Asset Price Volatility. Federal Reserve Bank of Chicago Working Paper WP-94-2.

Mishkin, F., (1990). The Information in the Longer Maturity Term Structure about Future Inflation. *Quarterly Journal of Economics*, **105**: 815-828.

Pencavel, J., (1986). Labor Supply of Men: A Survey. *Handbook of Labor Economics*, (eds.) O. Ashenfelter and R. Layard. Amsterdam: North-Holland.

Rotemberg, J. and Woodford, M., (1995). Dynamic General Equilibrium Models with Imperfectly Competitive Product Markets. *Frontiers of Business Cycle*, (ed.) Thomas F. Cooley. Princeton: Princeton University Press.

Schroder, M., (1995). Optimal Portfolio Selection with Fixed Transaction Costs. Working Paper, State University of New York at Buffalo.

Sims, C., (1980). Macroeconomics and Reality. *Econometrica*, **48**: 1-48.

Sims, C., (1986). Are Forecasting Models Usable for Policy Analysis? Federal Reserve Bank of Minneapolis *Quarterly Review*, **10**: 2-16.

Sims, C., (1992). Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy. *European Economic Review*, **36**: 975-1000.

Sims, C. and Zha, T., (1995a). Does Monetary Policy Generate Recessions? Working Paper, Yale University.

Sims, C. and Zha, T., (1995b). Error Bands for Impulse Responses. Federal Reserve Bank of Atlanta Working Paper No. 95-6.

Watson, M., (1994). Vectorautoregressions and Cointegration. *Handbook of Econometrics*, (eds.) R. Engle and D. McFadden. Amsterdam: North-Holland.