Permanent versus Transitory Changes in Money Growth and the Effects on the Term Structure

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Abstract

This paper uses a cointegrated structural VAR to decompose the monetary base growth rate into permanent and transitory components. The effects of changes in these components on both long-term and short-term interest rates are analyzed.
INTRODUCTION

The term structure of interest, or the yield curve as it is also called, has long been considered an important predictor of the health of the U.S. economy. Figure 1 displays the spread or difference between yields on the ten-year U.S. Treasury note and the three-month U.S. Treasury bill from January of 1960 to December of 1994, which will be called the term structure for simplicity. The shaded periods represent the NBER-dated recessions. Of the six recessions that have occurred over this interval in U.S. economic history, four are associated with inverted yield curves, i.e., the short-term interest rate is higher than the long-term rate. The remaining two recessionary episodes are coincident with a substantial flattening in the term structure, if not an outright inversion of it. Of the five times that the yield curve was inverted over this period, four were consistent with recession. Given this strong association between the yield curve and economic output in the U.S. economy, it seems prudent to gain a better understanding of how monetary policy affects the term structure of interest rates.

In this paper I analyze the effects of both permanent and temporary changes in the growth rate of the monetary base on nominal interest rates at both the short-term and long-term horizons. The methodology used is similar to the structural vector autoregression (SVAR) methods employed by Blanchard and Quah (1989), Gali (1992), and others. It differs in that existence of cointegration among the variables is explicitly accounted for and the restrictions implied by it are used to help identify a structural model. The main empirical result is that permanent increases in the growth rate of the monetary base temporarily flatten the yield curve, and depending upon the initial spread between long-term and short-term interest rates, may even invert the yield curve. This result is somewhat surprising but does have a plausible explanation. In the long-run, all permanent money growth changes effect both short-term and long-
term interest rates equally so that the long-run effect on the yield curve is to leave it unchanged. Temporary increases in the growth rate of the monetary base have the more predictable liquidity effect where short-term rates decline temporarily with little change in long-term rates. The effect is to steepen the yield curve or increase the spread between long-term and short-term rates. The effects of a transitory inflation shock are also produced as a by-product of the specific restriction used to identify the structural innovations. The effects of a temporary increase in inflation expectations raises long-term rates more than short-term rates on average inducing a steepening of the term structure for a short period of time.

The next section discusses the theoretical relationships between money growth and nominal interest rates and various term-to-maturities. Section 3 details the identification of the structural model and the related econometric techniques. Section 4 presents and discusses the empirical results and section 5 concludes.

THE LINK BETWEEN MONETARY POLICY AND INTEREST RATES

The relationship between money growth and interest rates is most easily demonstrated by using two simple but important economic theories. The first is the Fisher effect, named after Irving Fisher (1930), which relates the nominal and real interest rates via the expected loss in the purchasing power of money, usually proxied by general price inflation as in equation (1).

\[
(1 + i_t^e) = (1 + r_t)(1 + E_t \pi_{t+\ell})
\]

The long-run (super-) neutrality of money implies that the expected inflation term in equation (1), \(E_t \pi_{t+\ell}\), can be replaced by the expected money growth rate (see Lucas, 1980). If the ex-ante real rate is constant, as suggested by Fama (1975), then changes in the expected growth rate of money will lead to one-for-one changes in
the short-term nominal interest rate, \( i_t \). Crowder and Hoffman (1996) demonstrate that while nominal rates of interest do move in tandem with expected inflation as suggested by theory, it can take up to several years for complete adjustment to occur. This slow adjustment of inflation expectations leaves considerable room for short-run real interest rate effects implying that changes in money growth can affect real rates of interest in the short to medium term.

The linkage between money growth and long-term interest rates is somewhat more circuitous. The expectations theory of the term structure implies that long-term interest rates should be equal to the expected yield of rolling over the requisite number of short-term rates to achieve the same term-to-maturity. Equation (2) gives the standard formulation where subscripts denote the period that the asset is purchased and superscripts denote the term-to-maturity and \( E_t \) is the expectations operator conditioned on information available in period \( t \).

\[
(1 + i_t^t) = E_t \prod_{j=1}^{\ell} (1 + i_j^j)
\]  

Equation (2) can explain the shape of the yield curve by noting that if future short-term rates are expected to rise, then the current long-term rate will exceed the current short-term rate and we would observe the classic upward sloping yield curve. If future short-term rates are expected to decline, then the yield curve will be inverted or downward sloping. If short-term rates are anticipated to remain constant, then the yield curve will be flat. Two factors that have been theorized to drive a wedge between expected future short-term rates and current long-term rates are the term or liquidity premium and the inflation risk premium. The term or liquidity premium arises when agents investment horizon is shorter than the term-to-maturity on a particular asset. In order to be induced to hold such assets they must be compensated for the greater risk with a higher yield. The inflation risk premium arises due to the increased uncertainty about the future purchasing power of money as the term-to-maturity
increases. Neither of these addendums to the expectations theory overturns the major implications of it. But the addition of risk premia can help explain departures from the pure expectations theory of the term structure.

The effect of changes in money growth on long-term rates is determined by the effect of changes in money growth on expected future short-term interest rates. Two possibilities exist; if the change in the growth rate of money is purely permanent, such as a permanent increase in the growth rate from 1% per annum to 2%, then expected future short-term rates will be higher by 1% and the current long-term rate will also rise. But if the change in the growth rate is purely temporary, then expected future short-term interest rates will remain unchanged implying that current long-term rates will also not change. In the short-run, however, the excess liquidity created by the temporary increase in money growth may lead to increased purchases of assets which would create a temporary decline in the rate of return on the assets. This is the so-called "liquidity effect". Working against the liquidity effect on long-term interest rates are the term and inflation risk premia.

In the next section I outline the procedure used to estimate a dynamic structural model of money growth and short-term and long-term interest rates. The time series are assumed to be individually non-stationary, an assumption that will be supported later, but collectively to be represented by only one non-stationary component. The assumption of a single common stochastic trend is implied by log-linearized versions of equations (1) and (2) under the further assumptions that expectations are weakly rational and the (ex-ante) real interest rate is a stationary series.¹ The existence

¹Weakly rational expectations allow for temporary deviations from pure rational implications. For example, rational expectations requires that the expectations error be serially uncorrelated. Weakly rational expectations simply requires that expectations errors not be persistent for long periods of time. The assumption of a stationary real interest rate is also supported theoretically. See Crowder and Hoffman (1996) for a perspicuous discussion.
of the common trend implies that the proper reduced form dynamic specification is
the error correction model (ECM) of Engle and Granger (1987). The log-linearized
versions of (1) and (2) imply that money growth should be cointegrated with the
short-term interest rate with a cointegration vector of $[1, -1]'$ and the short-term
interest rate should be cointegrated with the long-term rate also with cointegrating
vector $[1, -1]'$. These restrictions are tested in section 4.

IDENTIFICATION OF A STRUCTURAL MODEL

The identification of a structural model from the dynamic reduced form has been
extensively discussed since Sims (1980) first proposed the use of vector autoregressions
(VARs) in analyzing the dynamic relationship among time series variables. The
issues surrounding structural identification can be understood by specifying a reduced
form VAR in (3),

$$\Phi(L)X_t = \mu + \varepsilon_t$$

where $X_t$ is a $p \times 1$ vector of stationary time series variables, $\Phi(L)$ is a matrix poly-
nomial in the lag operator where $\Phi_0 = I$, $\mu$ is a vector of deterministic components
(usually a constant) and $\varepsilon_t$ is the reduced form error with covariance matrix $\Omega$. A
corresponding structural model is assumed to exist as in (4),

$$A(L)X_t = \delta + \nu_t$$

where $E[\nu'_t \nu_t] = I$. The relationship between the structural and reduced form param-
eters is $\Phi(L) = A_0^{-1}A(L)$, $\mu = A_0^{-1}\delta$, and $A_0\Omega A_0' = I$. Identification of the structural
model from the reduced form estimates is simply an exercise in specifying $A_0$ in such

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2The relevant citations include, but are not limited to, Bernanke (1986), Sims (1986), Shapiro

3Thus, $X_t$ could be the first differences of I(1) variables.
a way that a unique correspondence exists between the reduced form and structural parameters. In general, this would require \( p^2 \) independent restrictions on (4). Arbitrary normalization of each reduced form VAR equation provides \( p \) restrictions. The assumption of structural error independence provides another \( p(p-1)/2 \) restrictions. Exact identification of the structural model then depends upon \( p(p-1)/2 \) additional restrictions to be imposed by the econometrician. These restrictions take the form of restrictions on the \( A_0 \) matrix.

Sims (1980) original specification made \( A_0 \) a lower triangular matrix implying a contemporaneous recursive structure to the model. Bernanke (1986) and Sims (1986) independently suggested applying non-recursive exclusion and/or general restrictions on \( A_0 \) that were more justifiable based upon economic reasoning. Shapiro and Watson (1988) and Blanchard and Quah (1989) used long-run restrictions to identify the structural model from the reduced form estimates. This is achieved by noting that the VARs in (3) and (4) have invertible Wold moving average representations (MAR). The reduced form MAR is given by,

\[
X_t = \zeta + C(L)\varepsilon_t
\]

(5)

where \( C(L) = \Phi(L)^{-1} \) and \( \zeta = \Phi(L)^{-1}\mu \). The corresponding structural MAR is related to the reduced form MAR by noting the \( D(L) \), the structural MA matrix polynomial, is equal to \( C(L)A_0^{-1} \). Furthermore, the long run structural total impact matrix is related to the reduced form total impact matrix by \( D(1) = C(1)A_0^{-1} \). General restrictions on the \( D(1) \) matrix can be imposed by a suitable choice for \( A_0 \). These are then interpreted as long run as opposed to contemporaneous restrictions on the interactions of the variables in \( X_t \). Blanchard and Quah (1989) argue that economic theory provides stronger implications for the long run interaction among macro aggregates than it does about the contemporaneous relationship between these variables. Such "neutrality" restrictions are more justifiable on theoretical grounds.
Note that for the long-run identification restrictions to be binding, some of the variables in $X_t$ must be the first differences of nonstationary variables. That is the $C(1)$ matrix must not vanish for these type of restrictions to be relevant. This issue is important since many important examples exist in macroeconomics and finance where the $C(1)$ matrix is not of full rank. This implies that only a subset of the innovations have a permanent impact on the variables. The other innovations have only transitory effects and cannot be identified from restrictions on $C(1)$. A relevant situation is when the variables of interest are cointegrated. The model used in this study sets $p = 3$ implying that the number of additional a priori restrictions needed to achieve exact identification is three, $p(p - 1)/2$. Assuming that the variables are I(1), the model also implies two cointegrating relations among the three variables. This means that the rank of $C(1)$ is one or that only one of the three structural innovations leaves permanent imprints on the data.\footnote{Strictly speaking, the condition that $C(1)$ has rank of one implies that either only one of the structural innovations has a permanent effect on the variables or that two or more of the structural innovations has identical permanent effects on the variables. This seems to be an unlikely scenario but cannot be completely ruled out.} Furthermore, $C(1)$ must satisfy the restrictions implied by cointegration\footnote{See Engle and Granger (1987) and Johansen (1991).}, namely that $\beta'C(1) = 0$ and $C(1)\alpha = 0$, where $\alpha$ and $\beta$ are $p \times 1$ full rank matrices such that $\alpha\beta' = \Phi(1)$ from (3). These imply two restrictions on $C(1)$ that help aid in the identification of $D(1)$.

There are two implications from the foregoing discussion. The first and most obvious is that the knowledge of cointegration rank and space, in general, will yield restrictions that reduce the number of a priori restrictions needed to achieve exact identification of the structural model. The second, and more subtle implication is that existence of cointegration among the variables reduces the set of possible restrictions that the econometrician can legitimately impose. For example, in the three variable case to be analyzed in the next section, there exists two cointegrating rela-
tions. Therefore, no long run neutrality restrictions can be imposed on two of the three innovations since they are already neutral, with respect to the variables in the system, by definition. That is, such a restriction is not binding nor independent of the cointegrating restrictions. In general then, the econometrician needs to combine both long run and contemporaneous restrictions to achieve exact identification of the structural model. Specifically, when there exists $r$ cointegrating relations among $p$ variables, there will be $k = p - r$ permanent innovations that can be identified via long run restrictions and $r$ transitory innovations that can only be identified using short run or contemporaneous restrictions. The restrictions on $C(1)$ implied by the cointegrating relations yields $kr$ independent restrictions leaving just $p(p - 1)/2 - kr$ independent \textit{a priori} restrictions needed to just identify the model. Of these $k(k - 1)/2$ should be long run restrictions and $r(r - 1)/2$ should be contemporaneous or short run restrictions.

\footnote{Crowder (1995a) demonstrates that Blanchard and Quah’s original model can be represented as a cointegrated system and that the long run neutrality restriction they impose is actually overidentifying. The test of the over-identifying restriction is not rejected in the Blanchard and Quah data but is rejected when an updated sample is used. This example demonstrates the special case where some of the variables in a system are $I(1)$ and others are $I(0)$. This system can (and probably should) be represented as a cointegrated system, albeit a trivial form of cointegration. Specifically, the result that the system is driven by a reduced number of (stochastic) trends must be accounted for in the identification stage.}

\footnote{Gali (1992) does this but ignores the cointegrating relationship between real M1, income and nominal interest rates and is thus still subject to the problem discussed. See Crowder, et al. (1995) for an analysis that properly accounts for the Fisher and equilibrium money demand relations to identify a structural model.}
EMPIRICAL RESULTS

Data

The data used in this study represent monthly observations beginning in January 1960 and ending in December 1994 on the growth rate of the seasonally adjusted U.S. monetary base adjusted for reserve requirements maintained by the Federal Reserve Bank of St. Louis and the interest rates on U.S. 90-day Treasury bills and 10-year Treasury notes.\footnote{In unreported estimates that are available upon request, I also included interest rates on 180-day, 360-day and 5-year U.S. Treasury securities as well as the Federal Funds rate with no appreciable difference in results. In addition, I analyzed the sensitivity of the results to alternative money definitions such as M1 and M2 and the adjusted base measure maintained by the Federal Reserve Board. The results are qualitatively unchanged from those presented in the paper.} Each series is plotted in figure 2, along with the spread between the 10-year note and 90-day bill rates.

Cointegration Analysis

Each of the three series appear to be non-stationary. A host of univariate unit root tests confirms that each of the series is integrated of order one.\footnote{The results are available upon request.} The estimation and testing of the long run equilibrium (cointegrating) relationship was done using the Johansen (1988) method.\footnote{The methodology is well known, so I will forgo a discussion here. The uninitiated reader is referred to the citations listed in the references.} Table 1 presents the summary statistics from the cointegration analysis. The specification of the deterministic components of the VAR is important in determining the limiting distribution of the cointegration test statistics (see Johansen, 1994). I begin with the most general specification that is relevant, i.e., a constant in the first-differenced VAR implying the existence of linear trends in the data that are eliminated by the cointegrating relationship. The hypothesis of...
no cointegration under this specification is denoted as $H_1(r)$. This specification is
tested against the restriction that there are no deterministic trends in the data only
a non-zero constant in the equilibrium relationship. The null hypothesis of no cointe-
gration under this specification is $H_1^*(r)$. The form of the test is discussed in Johansen
(1994). The calculated test statistic of 0.227 is asymptotically distributed as a $\chi^2(1)$
variate. The restriction cannot be rejected. The restriction that the constant in the
equilibrium relationship is equal to zero is also tested. The null hypothesis of no coin-
tegration is denoted by $H_0(r)$ in this specification of the deterministic components.
This yields a test statistic of 5.158 which is distributed as a $\chi^2(2)$ variate. This re-
striction cannot be rejected at the 5% level of significance. Therefore, the estimation
and testing of the existence of a single common stochastic trend, as implied by the
theory in section 2, is conducted under the restriction that there is no constant in
the cointegration vectors.

Moving Average Representation and Identification

From the results presented in table 1 it can be concluded that there is strong
evidence in favor of two cointegration vectors among the three time series. This means
that there is only one common trend among the three variables as predicted by the
theory in section 2. Furthermore, the normalized cointegration space is statistically
insignificantly different from that implied by theory as seen in table 2.

To analyze the dynamic effects of innovations in the structural errors, the MAR is
needed. The existence of cointegration implies that the $\Phi(L)$ matrix has a singularity
which complicates the inversion to yield the MAR. The MAR is computed by noting
that $C(1) = \beta_1(\alpha_1^*\Phi^*(1)\beta_1)^{-1}\alpha_1'$ (see Johansen, 1991), where $\beta'\beta_1 = 0$, $\alpha'\alpha_1 = 0$,
$\Phi^*(1) = \Phi_1^* + ... + \Phi_{k-1}^*$ and $\Phi^*(L) = (\Phi(L) - \Phi(1))(1 - L)^{-1}$. Johansen (1991) has
given $\alpha_1'$ the interpretation of the common trends, as in Stock and Watson (1988),
and $\beta_1(\alpha_1^*\Phi^*(1)\beta_1)^{-1}$ is interpreted as the factor loadings or how the common trends
are passed on to the variables in the system. Equation (1) implies that the stochastic trend in the short-term interest rate is due to the stochastic trend in (expected) inflation. Extending the logic, if the trend in inflation is due to the trend in money growth (see Crowder, 1996) then it must be the case that the trend in the short-term rate is also due to the trend in money growth. Furthermore, equation (2) implies that the stochastic trend in the long-term rate is due to the trend in the short-term rate. Ultimately then, the common stochastic trend in \( X_t = [m_t, i_t^s, i_t^L]' \) must be the money growth measure, \( m_t \). This implies that only the innovations in the money growth variable accumulate to form the common trend, i.e., \( \alpha_\perp = [1, 0, 0]' \). This restriction is testable within the MLE framework of Johansen by using a likelihood ratio (LR) test. The LR statistic associated with the hypothesis that \( \alpha_\perp = [1, 0, 0]' \) is 2.88 which is distributed as a \( \chi^2(2) \) variate. This restriction cannot be rejected at the 5% level of significance. Testing the restrictions that either \( \alpha_\perp = [0, 1, 0]' \) or \( \alpha_\perp = [0, 0, 1]' \) yields LR statistics of 10.86 and 13.57, respectively. These are also \( \chi^2(2) \) variates and thus can both be rejected at high levels of significance. The empirical results support all of the restrictions implied by the theoretical models in (1) and (2).

Identification of the structural model can proceed. Given that \( C(1) = \beta_\perp (\alpha_\perp \Phi^s(1)\beta_\perp)^{-1}\alpha_\perp' \) and that \( C(1) = A_0 D(1) \), restrictions on \( C(1) \) can be imposed either through restrictions on the common trends, \( \alpha_\perp' \), or the factor loadings, \( \beta_\perp (\alpha_\perp \Phi^s(1)\beta_\perp)^{-1} \). The structural total impact matrix \( D(1) \) can be written as \( \beta_\perp^0 (\alpha_\perp \Phi^s(1)\beta_\perp^0)^{-1}\alpha_\perp' \) where \( \beta_\perp^0 = A_0 \beta_\perp \). Thus the identifying restrictions can be imposed directly on \( \beta_\perp \).\(^{11}\) Since there is only one common trend or \( k = 1 \), no extra \textit{a priori} restrictions are needed.

\(^{11}\)The choice of \( \beta_\perp^0 \) must also satisfy the condition that the structural errors are independent. This is ensured when \( (\beta_\perp^0 \beta_\perp^*)^{-1} \beta_\perp^* C(1) \Omega C(1)' \beta_\perp^* (\beta_\perp^0 \beta_\perp^*)^{-1} \) is diagonal, where \( \beta_\perp^* = \beta_\perp^0 (\alpha_\perp \Phi^s(1)\beta_\perp^0)^{-1} \). Taking the Cholesky factor of this matrix and multiplying it by \( \beta_\perp^0 \) will achieve this independence requirement.
to achieve identification of the permanent innovations. But, since \( r = 2 \), i.e., there are two cointegrating vectors one extra restriction is necessary in order to separately identify the two transitory innovations. The existence of cointegration has supplied two of the three restrictions needed to exactly identify the structural model, aside from the assumption of structural error independence.

The restrictions used to identify the transitory innovations can be imposed by using the following procedure suggested by Warne (1993). Since the transitory innovations must be orthogonal to the permanent innovations and the covariance between the transitory and permanent errors is \( \alpha' \Omega \alpha_0 \), a suitable choice for \( A_0 \), the \( r \) rows of \( A_0 \) associated with the transitory innovations, will be \( \alpha_0 \Omega^{-1} \), where \( \alpha_0 \) is any space spanned by the columns of \( \alpha \) so that \( \alpha' \alpha_\perp = 0 \). To ensure that the transitory innovations are orthogonal to each other, i.e., \( \alpha_0 \Omega^{-1} \alpha_0 \) is diagonal, one may use any of the standard orthogonalization procedures such as Sims (1980), a Cholesky decomposition, or Bernanke (1986), a structural non-recursive decomposition. In the application here \( r = 2 \) so that the exact form of orthogonalization is irrelevant. I choose to orthogonalize the transitory innovations by taking a Cholesky decomposition of \( \alpha_0 \Omega^{-1} \alpha_0 \).

The variables in each system are ordered as \( X_t = [m_t, i_t^S, i_t^L] \), where \( i^S \) is the short-term interest rate and \( i^L \) is the long-term interest rate. Since the transitory innovations have, by definition, no long-run effect on any of the variables, one cannot use long-run neutrality restrictions to achieve identification as in Blanchard and Quah (1989). The restriction needs to be on the contemporaneous impact matrix. The single extra identification restriction that I impose is that one of the two transitory innovations has no contemporaneous effect on the growth rate of the monetary base. This innovation can then be thought of as either a real interest rate innovation or an inflation expectations innovation. In either case, nominal money growth will respond to shocks to this innovation with a one period lag. The other transitory innovation
then can be given the interpretation as a temporary money growth shock, while the permanent innovation can be interpreted as a permanent shock in money growth.

Impulse Response Analysis

Figure 3 plots the accumulated impulse response functions (IRFs) for the each variable due to a permanent structural shock to money growth. The shock is normalized to have a long-run impact on money growth of 1%. This implies that the long-run impact on both short-term and long-term interest rates is also 1%. The long-run impact on the term structur is, however, zero due to the stationarity of the spread between long-term and short-term interest rates. The short-run effects of a permanent money growth shock are quite interesting. The immediate effect on the short-term interest rate is quite large, essentially causing it to jump immediately to its new long-run value. The effect on the long-term rate is less drastic. It takes fully 24 months for the full impact to occur. Over the period of adjustment the yield curve or term structure is severely flattened or inverted. This adjustment process suggests that agents do not fully anticipate the permanence of the money growth shock and that some sort of learning is going on.

Figure 4 presents the IRFs from a transitory money growth shock. The familiar pattern of the liquidity effect is evident in the short-term rate response. There is no liquidity effect on the long-term rates which is consistent with the notion that since there is no change in the long-run expected inflation and no change in the expected future short-term rates there is no change in the current long-term rate. The yield curve becomes much steeper after a transitory money growth shock. From figure 1 it can be seen that a steeper yield curve is associated with economic expansion. The periods in which the spread between long-term rates and short-term rates is the greatest are also periods of economic expansion.

Figure 5 displays the IRFs from the transitory inflation expectations shock. The
innovation causes both rates to rise in the short-run and money to decline, presumably a response by the monetary authority. The net effect on the yield curve is to steepen it, but this effect is not statistically significant.

**Variance Decomposition**

Table 3 presents the decomposition of the model forecast errors for the term structure. These results highlight the importance that transitory money growth shocks play in explaining the behavior of the term structure over the post-war period. Approximately 50% of the forecast error variance of the term structure is explained by the transitory money growth shock for up to twelve months. The other 50% is explained by permanent money growth shocks. The transitory inflation expectations shocks play a negligible role in explaining the term structure behavior at all forecast horizons. Even at very long forecast horizons the transitory money growth shock explains a significant portion of the variance in the term structure.

**Historical Decompositions**

In this section, the importance of each structural innovation in explaining the behavior of the term structure over particular historical episodes is discussed. Figure 6 presents the historical dynamic forecast error of the term structure from the estimated model, with recessions denoted by the shaded regions. Two things can be gleaned from figure 6; first, the statistical model does a relatively good job of forecast the term structure, i.e., the spread between long-term and short-term interest rates; second, the model’s forecast error variance increases appreciably during the recessions of 1973-75, 1980, 1981-82, and, to a lesser extent, 1969-70. This implies that these changes in the term structure that occurred around the periods of significant economic downturn are not predicted by the model. The behavior of the yield curve in these episodes is
largely due to the unforecastable structural innovations.

Figures 7-9 depict the portions of the total dynamic forecast error in figure 6 that are attributable to each of the structural innovations. Figure 9 displays the fraction of the dynamic forecast error that is due to the inflation expectations shock. As suggested by the variance decompositions, the transitory innovation in expected inflation has little explanatory power for the behavior of the term structure in any period. Figure 7 shows the part of the historical forecast error due to the permanent money growth innovation. This shock explains a significant amount of the unexplained behavior in the term structure in the 1979-80 and 1981-82 recessions. It also captures movements at the end of both the 1973-75 and 1990-91 recessions. Figure 8 presents the historical decomposition of the term structure forecast error attributed to the transitory money shock. This innovation explains a great deal of the movement in the forecast error during all of the recessions, implying that much of the unexplained behavior of the term structure during and around the time of recessions is due to temporary changes in the growth rate of the monetary base that are not anticipated.

CONCLUSIONS

This paper has used a simple cointegrated structural VAR to analyze the effects of shocks to money growth and expected inflation on the term structure of interest rates. The results suggest that while the model does a fairly good job of forecasting the changes in the yield curve, recessions are still not very predictable.
REFERENCES


Table 1. Johansen Trace Tests.

### Eigenvalues and Trace Statistics: $H_1(r)$

<table>
<thead>
<tr>
<th>Test</th>
<th>$r = 0$</th>
<th>$r \leq 1$</th>
<th>$r \leq 2$</th>
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<tr>
<td>$\hat{\lambda}$</td>
<td>0.058</td>
<td>0.042</td>
<td>0.013</td>
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<tr>
<td>$LR_{trace}$</td>
<td>46.88</td>
<td>22.48</td>
<td>5.17</td>
</tr>
</tbody>
</table>

5% C.V. \begin{align*} 29.68 & 15.41 \ 5.76 \end{align*}

### Eigenvalues and Trace Statistics: $H_0(r)$

<table>
<thead>
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<th>$r \leq 1$</th>
<th>$r \leq 2$</th>
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<tr>
<td>$\hat{\lambda}$</td>
<td>0.048</td>
<td>0.040</td>
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<tr>
<td>$LR_{trace}$</td>
<td>36.72</td>
<td>16.49</td>
<td>0.01</td>
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</tbody>
</table>

5% C.V. \begin{align*} 26.64 & 12.53 \ 3.84 \end{align*}

Note: Critical values are taken from Osterwald–Lenum (1992).
Table 2. Estimates of the Normalized Cointegration Space.

\[ x_t = \begin{bmatrix} m_t & i_t^S & i_t^L \end{bmatrix} \]

Normalized Estimate of Cointegration Space

\[ \hat{\beta} = \begin{bmatrix} 1.00 & -1.04(0.11) & 0.00 \\ 1.00 & 0.00 & -1.13(0.10) \end{bmatrix} \]

Note: Numbers in parentheses are asymptotic standard errors.

Table 3. Forecast error variance decomposition.

<table>
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<th>Series</th>
<th>shock</th>
<th>1</th>
<th>4</th>
<th>8</th>
<th>12</th>
<th>48</th>
<th>120</th>
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<td>Term Structure</td>
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<td>45.91</td>
<td>47.96</td>
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<td></td>
<td>tmp1</td>
<td>57.17</td>
<td>52.58</td>
<td>49.33</td>
<td>43.41</td>
<td>36.38</td>
<td>36.33</td>
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<td></td>
<td>tmp2</td>
<td>5.73</td>
<td>1.51</td>
<td>2.72</td>
<td>4.99</td>
<td>10.00</td>
<td>10.10</td>
<td>10.00</td>
</tr>
</tbody>
</table>

Note: The restriction used to separate the effects of the two orthogonal transitory innovations is that tmp₂, the transitory inflation expectations shock, as no contemporaneous effect on the money growth rate.
Fig. 1. Ten-Year Treasury Note Interest Rate Minus Three-Month Treasury Bill Interest Rate

Fig. 2. Data Used in the Analysis
Fig. 3. Impulse Responses from a Permanent Money Growth Shock

Fig. 4. Impulse Responses Due to a Transitory Money Growth Shock
FIG. 5. Impulse Responses Due to an Inflation Shock

FIG. 6. One-Step-Ahead Dynamic Historical Forecast Error
FIG. 7. Contribution of the Historical Forecast Error Due to the Permanent Money Growth Innovation

FIG. 8. Contribution of the Historical Forecast Error Due to the Transitory Money Growth Innovation
FIG. 9. Contribution of the Historical Forecast Error Due to the Inflation Innovation