News Shocks and the Slope of the Term Structure of Interest Rates

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Abstract

We provide a new structural interpretation of the relationship between the slope of the term structure of interest rates and macroeconomic fundamentals. We first adopt an agnostic identification approach that allows us to identify the shocks that explain most of the movements in the slope. We find that one shock can explain the majority of all unpredictable slope innovations over a 10-year forecast horizon. Impulse response functions lead us to interpret this shock as a news shock about future productivity. We confirm this interpretation by formally identifying such a news shock as in Barsky and Sims (2009) and Sims (2009). We then assess to what extent a New Keynesian DSGE model is capable of generating the observed slope responses to a news shock. While the model provides valuable information on how TFP news shocks transmit through the economy to the slope of the term structure, we find that the model falls short of matching quantitatively the joint macro and term structure dynamics observed in the data.

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1 Introduction

The slope of the term structure – commonly defined as the spread between the yield on a long-term treasury bond and a short-term bill rate – has drawn the attention of many separate literatures. In forecasting, it is well established that the slope provides valuable predictive content for future economic activity (e.g. Estrella and Hardouvelis, 1991).\(^1\) In finance, there is a large literature trying to explain both the average size and the time-variation of the slope with either latent factor no-arbitrage models (e.g. Duffie and Kan, 1996) or consumption-based asset pricing models (e.g. Piazzesi and Schneider, 2006).\(^2\) In macroeconomics, the slope of the term structure plays a central role for the transmission of monetary policy (e.g. Clarida, Gali and Gertler 1999). In recent years, different papers have attempted to bridge the gap between these literatures. While these papers have uncovered strong linkages between term structure and macroeconomic dynamics (e.g. Ang and Piazzesi, 2003; Diebold, Rudebusch and Aruoba, 2006), important questions remain unanswered. In particular, what are the fundamental sources of movements in the term structure slope? Do these fundamentals look like typical macroeconomic shocks? Can modern dynamic stochastic general equilibrium (DSGE) models replicate the observed slope responses to these shocks?

In this paper we provide answers to these questions. We first use a macro-finance VAR to show that over 60% of movements in the slope are due to news shocks about future innovations to total factor productivity (TFP). A key driver of this result is the endogenous response of monetary policy. After a positive news shock, the Federal Funds rate, and with it the short-end of the term structure, falls. Since the reaction of the long-end of the term structure is small, the slope increases and only gradually returns to its initial value. The shock we identify provides a unified explanation for a number of stylized facts about the term structure: (i) variations in the slope are primarily due to fluctuations in the short-end of the term structure; (ii) steep yield curves (i.e. large slopes) generally predict future economic growth; and (iii) systematic monetary policy plays an important role for the linkage between macroeconomic and term structure dynamics. We then assess to what extent a modern medium-scale DSGE model is capable of generating the observed slope responses to a news shock. While the model provides valuable insights into the linkages between the macroeconomy and the term structure, we find that the model falls short of matching quantitatively the joint macro and term structure dynamics to a TFP news shock.

\(^1\)See Ang, Piazzesi and Wei (2006) for a recent application and an extensive review of the literature.

\(^2\)Other important latent factor no-arbitrage contributions include Knez, Litterman and Scheinkman (1994) and Dai and Singleton (2000). Recent consumption-based contributions are Wachter (2006) and Bansal and Shaliastovich (2007).
observed in the data. In our view, this failure represents an important challenge for modern macroeconomic models.

Our interest in uncovering the sources of term structure fluctuations is motivated by the recent macro-finance literature that has taken the first step of linking simple atheoretical term structure factor models with macroeconomic factors. In this paper we take the next step of uncovering the fundamental shocks and structure that propagates these shocks between macroeconomic and financial variables. To do so, we adopt a novel identification strategy. As in existing papers, we start by combining term structure variables with prominent macro aggregates in a VAR. But instead of postulating a particular type of shock and then analyzing its effects, our strategy consists of first uncovering (in a statistical sense) the shocks that are quantitatively most important for the slope of the term structure. We then provide an economic interpretation of these shocks. To do so, we apply a methodology developed by Uhlig (2003) to extract the exogenous shocks that explain as much as possible of the Forecast Error Variance (FEV) of a target variable in the VAR, which in our case is the slope. We then interpret these shocks by analyzing the impulse response functions (IRFs) of the different variables in the VAR and contrasting them with the theoretical implications of typical macroeconomic shocks.

Nothing in our approach requires that a small number of shocks accounts for a large fraction of slope variations or that these shocks have an appealing interpretation. Yet, when applying our empirical strategy to the 1959-2005 period, we find that one single shock can account for 70% to 90% of all unpredictable fluctuations in the term structure slope over a 10-year horizon. Furthermore, we find that this slope shock closely resembles a news shock about future innovations in TFP as proposed in Beaudry and Portier (2006), Jaimovich and Rebelo (2009) or more recently Barsky and Sims (2009) and Sims (2009). Specifically, TFP and consumption barely move on impact of the shock but gradually increase to a new permanent level thereafter. At the same time, inflation and the Federal Funds rate drop sharply and remain below their initial level for more than 2 years. The gradual but permanent long-run reaction of TFP and consumption together with the inverse reaction of both inflation and the Federal Funds rate rules out alternative interpretations of the slope shock such as exogenous monetary policy shocks, demand shocks, marginal rate of substitution shocks or contemporaneous TFP shocks.

To investigate the TFP news shock interpretation more formally, we follow Barsky and Sims (2009) and Sims (2009) and identify a TFP news shock directly as the innovation that accounts for most of the FEV of TFP over a 10-year horizon but is orthogonal to contemporaneous TFP movements. Even though this identification procedure is completely
different from our slope shock identification, we find that the extracted TFP news shock is highly correlated with the slope shock; explains over 60% of slope movements at all forecast horizons; and generates almost identical IRFs. These results remain unchanged for a battery of robustness checks. We conclude that the main driver of fluctuations in the slope of the term structure is news about future innovations to TFP.

To shed more light on the transmission mechanism from TFP news shocks to the term structure, we decompose slope movements into a part due to the Expectations Hypothesis and a part due to variations in term premia. We find that term premia increase significantly on impact of a positive TFP news shock. This is consistent with the general statistical rejection of the Expectations Hypothesis in the finance literature (e.g. Campbell and Shiller 1991, Cochrane and Piazzesi 2005). At the same time, the Expectations Hypothesis remains empirically relevant: the negative response of the Federal Funds rate and thus the short-end of the term structure to the TFP news shock is much stronger than the reaction of the discounted sum of future expected short rates (i.e. the long rate under the Expectations Hypothesis). The resulting difference accounts for more than half of the total increase in the slope. Hence, the systematic response of monetary policy is an important channel through which TFP news shocks transmit to movements in the slope.

In the last part of the paper we evaluate the extent to which a medium-scale DSGE model along the lines of Smets and Wouters (2007) can account for term structure movements in response to a TFP news shock. As is well known, linear approximations of DSGE models with homoscedastic innovations imply constant term premia (i.e. the Expectations Hypothesis holds exactly). Incorporating time variation in term premia is not straightforward. One approach to achieve this is to work with non-linear versions of DSGE models. Unfortunately, papers adopting such a non-linear approach have typically found it hard to generate sizable and sufficiently variable term premia.\footnote{\cite{Donaldson2012} and Den Haan (1995) are among the first to document the inability of basic DSGE models to generate large and volatile term premia. More recently, Rudebusch and Swanson (2008, 2009) and Binsbergen et al. (2010) among others reexamine the same issue for DSGE models with long-memory habits or recursive preferences.} Additionally, estimating non-linear DSGE models of the size considered here is computationally very challenging.

We pursue an alternative approach in the spirit of the latent factor no-arbitrage literature. As in Wu (2006) and Bekaert, Cho and Moreno (2010), we first derive an affine formulation of the pricing kernel from our DSGE model. Following Ang and Piazzesi (2003) and many others we add time-varying risk as a linear function of state variables. Long bond yields can then be derived recursively as a function of expected future short yields and time variation in term premia. In order to maintain parsimony, we restrict risk to depend on two key
variables of our DSGE model: expected inflation and expected changes in the marginal utility of consumption. Both of these variables have been shown in the finance literature to be important factors for the term structure. In contrast to that literature, however, the evolution of the two variables and thus risk is fully dictated by the solution to our linearized DSGE model. Furthermore, the parsimony of our setup restricts time variation in term premia in response to news shocks to depend on only two free parameters.

We estimate the parameters of our model using a minimum-distance methodology that tries to match the IRFs of the VAR to a TFP news shock. Our results show that while the estimated model is relatively successful in generating the response of macro aggregates to the news shock, the performance with respect to the term structure is more mixed. The model is capable of generating a persistent drop of the Federal Funds rate on impact of the shock. Quantitatively, however, the drop in the Federal Funds rate is less than half of what we observe in the VAR. As a result, the slope as implied by the Expectations Hypothesis increases only modestly and remains well below its empirical counterpart. To compensate, the model estimates a large increase in the term premium, which in turn leads to a counterfactual increase of the long rate on impact of the news shock.

Despite its quantitative failure, the DSGE model offers important insights into how TFP news shocks transmit through the economy to the slope of the term structure. Specifically, our VAR analysis implies that a positive TFP news shock triggers a substantial drop in inflation to which the Fed reacts systematically by lowering the Federal funds rate. Three crucial ingredients are necessary for the DSGE model to generate these responses. First, as in Barsky and Sims (2009), the model requires forward-looking price setting and a high degree of wage rigidity. Second, costs of capital utilization and adjusting investment need to be small. The combination of these elements implies that inflation drops on impact of the news shock. Third, monetary policy must react strongly to both inflation and output growth but not the output gap so as to generate a large decrease in the Federal funds rate. These results illustrate that augmenting DSGE models with news shocks and term structure variables provides valuable information to discipline the description of monetary policy and the structure of DSGE models in general.

Our paper is related to a number of studies on the linkages between term structure dynamics and macroeconomic fluctuations. In an innovative study, Piazzesi (2005) shows how to use high frequency data to trace the effect of exogenous monetary shocks onto yield data. Her work provides important insights into the nature of how monetary shocks work their way into the yield curve. Evans and Marshall (2007) combine term structure and macroeconomic variables in a VAR and identify fundamental innovations from empirical data.
measures of standard macroeconomic shocks. While the identified shocks have important
effects on the level of the term structure, they do not provide a quantitative explanation
for the majority of slope movements. This result motivates our approach of first finding the
shocks that are quantitatively important for the slope, and then interpreting them. At the
same time, it is important to note that our approach does not rule out the possibility that
other shocks play a significant role in slope movements. Our news shock, while a dominant
driver of the slope, still leaves up to 40 percent of the variation unexplained.

The DSGE literature has also begun to investigate the linkages between various macroeco-
nomic shocks and the term structure. Wu (2006), Rudebusch and Wu (2008) and Bekaert,
Cho, Moreno (2010) combine basic New-Keynesian models with no-arbitrage term structure
models to investigate the role of various shocks on yields. Wu (2006) and Bekaert, Cho and
Moreno (2010) conclude that monetary policy shocks explain a large portion of movements
in the slope. This contrasts with De Graeve, Emiris and Wouters (2009) who use a larger
DSGE model with many shocks (but no news shocks) and find that monetary policy shocks
play a much smaller role for the slope. Instead, demand shocks, defined as innovations to
the intertemporal consumption Euler equation, explain up to 50 percent of movements in
the slope. We interpret our results with respect to these papers as follows. Wu (2006),
Rudebusch and Wu (2008) and Bekaert, Cho and Moreno (2010) use relatively small models
with few shocks. If these models are too stylized or the number of shocks is too small,
the estimation may attribute movements in the short rate (which mostly drive the slope)
to monetary shocks since this shock is simply the residual of an interest-rate rule. This is
consistent with the results of De Graeve, Emiris and Wouters (2009) who argue, in addition,
that term premia become quantitatively less important once expectations of future short
rates are formed based on a larger DSGE model. We interpret their Euler equation shock
that explains up to 50 percent of the slope as a measurement error left to be explained
rather than a structural shock. Our news shock, in comparison, is one with a clear economic
interpretation and provides a 'deep' structural explanation for slope movements. Our results
also suggest that variations in term premia remain an important source of term structure
movements.

The remainder of the paper proceeds as follows. Section 2 explains our empirical ap-
proach. Section 3 provides information about the data and VAR specification. Section 4
presents our empirical results. Section 5 examines the dynamics of term premia. Section 6
presents the term structure DSGE model and estimates the model conditional on TFP news
shocks. Section 7 concludes.
2 Identifying Structural Shocks: Two VAR Approaches

In this section we present two approaches to VAR identification. The first approach, proposed by Uhlig (2003), is purely statistical and extracts the largest 1 or 2 (or 3 or 4) shocks that explain the maximal amount of the forecast error variance (FEV) in a target variable, which in our case is the slope of the term structure. The second identification approach is motivated by a key result from the first identification: news about future TFP play an important role in explaining movements in the slope. To assess this interpretation formally, we follow Barsky and Sims (2009) and Sims (2009) who extend the FEV maximization approach of Uhlig (2003) by using TFP as the target variable and imposing the extra restriction that the identified shock is orthogonal to contemporaneous TFP.

2.1 Review of VAR basics

We begin by discussing the general issue of identifying shocks in a VAR framework. This issue is well-known but we present these results for completeness and because the notation is useful for understanding the ensuing identification strategy. Consider a reduced-form VAR of the form

$$Y_t = B_1 Y_{t-1} + B_2 Y_{t-2} + \ldots + B_p Y_{t-p} + u_t,$$  \(1\)

where \(Y_t\) is a \(m \times 1\) vector of variables observed at time \(t\); and \(u_t\) is a \(m \times 1\) vector of one-step-ahead prediction errors with variance-covariance matrix \(E[u_t u_t'] = \Sigma\). Constant terms are dropped to save on notation. The objective is to impose restrictions on equation (1) to identify structural shocks; i.e. innovations that are mutually orthogonal to each other. Identifying all \(m\) shocks in the VAR requires a minimum of \(m(m-1)/2\) restrictions. However, it is well-known that one can instead place restrictions to identify fewer than \(m\) shocks.

In order to more clearly see the identification issue it is useful to rewrite equation (1). Under the assumption that \(Y_t\) is covariance-stationary, we can invert this VAR to express it as a moving average process

$$Y_t = [B(L)]^{-1} u_t = C(L) u_t,$$  \(2\)

where \(B(L) \equiv I - B_1 L - \ldots - B_p L^p\), and \(C(L) \equiv I + C_1 L + C_2 L^2 + \ldots\)

This moving average representation is of course the impulse response function for the VAR. Identification of the structural shocks amounts to decomposing the vector of prediction errors \(u_t\) into \(m\) mutually orthogonal innovations \(v_t\) with normalized variance-covariance matrix \(E[v_t v_t'] = I\). In other words, in identifying VAR shocks we are trying to find a
mapping $A$ between the reduced-form and structural shocks (i.e. $u_t = Av_t$).

In this mapping the $i$-th column of the $m \times m$ matrix $A$ describes the contemporaneous effect of the $i$-th innovation in the structural shock vector $v_t$ on the different variables in $Y_t$. By definition, $A$ needs to satisfy $\Sigma = E[Av_tv_t' A'] = AA'$. This restriction, however, is not sufficient to identify $A$ because for any matrix $A$, there exists some other matrix $\tilde{A}$ that satisfies the restriction that the covariance matrix be respected. This alternative matrix provides a different map from $u_t$ into $\tilde{v}_t$; i.e. $u_t = \tilde{A}\tilde{v}_t$. Thus, the set of statistically valid 'structural' identifications of the VAR is quite large. To choose which identification restriction to use, one then typically uses some sort of economic theory. One prominent example is to use the Cholesky decomposition to restrict $A$ to be lower triangular. Economically, this amounts to ordering the variables in the VAR in terms of the timing with which variables can respond to various structural shocks.

2.2 Extracting the Most Important Shocks

An alternative to the traditional approach of placing economic restrictions to identify a shock and then checking to see if the shock is important is to move in the reverse direction, as proposed by Uhlig (2003). This approach to identification is purely statistical and consists of finding the innovation(s) that explain(s) as much as possible of the FEV of some variable in $Y_t$ over a chosen horizon $k$ to $\bar{k}$. One then tries to provide an economic interpretation of the shock (conditional on it being important) by studying the full set of IRFs. As part of this procedure one learns how many shocks are needed to explain a given variable. That is, do we need a DSGE model with many shocks, or is a more parsimonious model able to explain a given time series?

More formally, Uhlig’s procedure searches for the $n$ largest shocks to explain the FEV of one variable in the VAR. Thus we need to find the $m \times n$ submatrix $A_1$ for the $n$ most important innovations in $v_t$ such that $A = [A_1 \ A_2]$ with $AA' = \Sigma$ for some $m \times (m-n)$ submatrix $A_2$. Given an initial decomposition $\tilde{A}$, this amounts to computing $A_1 = \tilde{A}Q_1$ where $Q_1$ is the $m \times n$ partition $Q_1$ of $Q = [Q_1 \ Q_2]$ that satisfies our statistical criteria.

To find $Q_1$, we let $\tilde{A}\tilde{A}' = \Sigma$ be the Cholesky decomposition of the reduced for VAR covariance matrix. We then define the impulse responses $\tilde{R}(L)$ associated with the innovations

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4To see this, consider an orthogonal matrix $Q$ with $QQ' = I$ and define $A = \tilde{A}Q$ and $Qv_t = \tilde{v}_t$. Then, $\Sigma = E[AQv_tv_t' A'\tilde{A}] = E[\tilde{A}\tilde{v}_t\tilde{v}_t' A'] = \tilde{A}\tilde{A}'$ because $E[\tilde{v}_t\tilde{v}_t'] = QE[v_tv_t']Q' = QQ' = I$.

5Any other triangular factorization would do as well, but the Cholesky is particularly easy to implement.
\( \tilde{u}_t \) identified by this decomposition as

\[
\tilde{R}(L) = C(L)\tilde{A} = \tilde{R}_0 + \tilde{R}_1 L + \ldots,
\]

with \( \tilde{R}_0 = \tilde{A} \). The impulse responses associated with the targeted innovations \( v_t \) are thus given by

\[
R(L) = C(L)\tilde{A}Q = \tilde{R}(L)Q.
\]

The \( k \)-step ahead forecast error of \( Y_{t+k} \) is then given by

\[
u_{t+k}(k) = Y_{t+k} - E_{t-1}[Y_{t+k}] = \sum_{l=0}^{k} \tilde{R}_l Q v_{t+k-l},
\]

and its variance-covariance matrix is given by

\[
\Sigma(k) = \sum_{l=0}^{k} \left( \tilde{R}_l[q_1 q_2 \ldots q_m] \left( \tilde{R}_l[q_1 q_2 \ldots q_m] \right)' \right)
\]

\[
= \sum_{j=0}^{m} \sum_{l=0}^{k} (\tilde{R}_l q_j) (\tilde{R}_l q_j)'
\]

where \( q_i, i = 1 \ldots m \) are the \( m \times 1 \) column vector partitions of \( Q \). The term \( \sum_{l=0}^{k} (\tilde{R}_l q_j) (\tilde{R}_l q_j)' \) thus describes the contribution of the \( j \)-th orthogonal shock to the variance-covariance matrix \( \Sigma(k) \) of the \( k \)-step ahead forecast error \( u_{t+k}(k) \). This division into \( m \) parts is possible because the \( v_t \) are \( iid \) innovations and the columns \( q_i, i = 1 \ldots m \) are orthogonal.

Our objective is to find the innovation(s) that explain(s) as much as possible of the sum of the \( k \)-step ahead forecast error variance of the \( i \)-th variable in \( Y \) over some horizon \( k \leq k \leq \bar{k} \)

\[
\sigma_i^2(k, \bar{k}) = \sum_{k=k}^{\bar{k}} \Sigma(k)_{ii}.
\]

Formally, to identify the most important innovation, we want to find the orthogonal vector
with length 1 (i.e. \( q_1'q_1 = 1 \)) that maximizes\(^6\)

\[
\sigma_i^2(k, \bar{k}; q_1) = \sum_{k=k}^{\bar{k}} \sum_{l=0}^{k} \left[ \tilde{R}_l q_1 (\tilde{R}_l q_1)' \right]_{ii} \\
= \sum_{k=k}^{\bar{k}} \sum_{l=0}^{k} \text{trace} \left[ E_{(ii)} (\tilde{R}_l q_1)(\tilde{R}_l q_1)' \right] \\
= q_1' S q_1,
\]

with

\[
S \equiv \sum_{k=k}^{\bar{k}} \sum_{l=0}^{k} \tilde{R}_l' E_{(ii)} \tilde{R}_l,
\]

where \( E_{(ii)} \) is a matrix with zeros everywhere except for the \( i, i \)-th position; and where the definition of \( S \) takes advantage of the fact that \( \text{trace} \left[ E_{(ii)} (\tilde{R}_l q_1)(\tilde{R}_l q_1)' \right] = \text{trace}((\tilde{R}_l q_1)' E_{(ii)} (\tilde{R}_l q_1)) \).

This maximization problem can thus be written as a Lagrangian

\[
L = q_1' S q_1 - \lambda(q_1'q_1 - 1)
\]

with first-order condition

\[
S q_1 = \lambda q_1.
\]

Inspection of this solution reveals that this is simply the definition of an eigenvalue decomposition, with \( q_1 \) being the eigenvector of \( S \) that corresponds to the eigenvalue \( \lambda \). Furthermore, since \( q_1'q_1 = 1 \), we can rewrite the first-order condition as \( \lambda = q_1' S q_1 = \sigma_i^2(k, \bar{k}; q_1) \). The partition \( q_1 \) that maximizes the variance is therefore the eigenvector associated with the largest eigenvalue \( \lambda \); i.e. \( q_1 \) is the first principal component of \( S \). Likewise, \( q_2 \) is the second principal component and so forth for all the \( n \) components of \( Q_1 \) that we want to extract.

The submatrix \( A_1 \) that we seek is then

\[
A_1 = \tilde{A} Q_1.
\]

### 2.3 Identifying News Shocks

Following Barsky and Sims (2009) and Sims (2009) (hereafter BSS), the shock that we seek to identify is news about future TFP. In their procedure, TFP is placed in a VAR with a selection of other macroeconomic variables. The assumption underlying the identification

\(^6\)For notational convenience, we order this vector first in \( Q \).
procedure is that TFP is an exogenous process. Therefore shocks to other variables in the system, such as monetary policy shocks, will not impact TFP at any horizon. Furthermore, it is assumed that TFP is driven by two shocks. One is an unforecastable shock to current TFP, while the second is a shock that represents news about future TFP.

The BSS identification approach extends Uhlig’s (2003) approach with additional restrictions. To implement the procedure we choose TFP as the variable in the VAR for which we would like extract the shocks to maximize the amount of the FEV explained. Further, the number of shocks is restricted to two (i.e. \( n = 2 \) in section 2.2). We then impose the additional restriction on the Lagrangian in (4) that the news shock must have zero impact on TFP contemporaneously.\(^7\) In other words, a news shock is defined as the innovation that explains most of future movements in TFP but nothing of current TFP. The other shock is then necessarily a contemporaneous TFP shock and can be identified by a Cholesky decomposition with TFP ordered first in the VAR.\(^8\)

3 Data and VAR Estimation Procedure

The VAR we estimate combines term structure and macroeconomic variables. For the term structure data we use two time series. The first is the Federal Funds rate. The second is the term spread which is measured as the difference between the 60-month Fama-Bliss unsmoothed zero-coupon yield from the CRSP government bonds files and the Federal Funds rate. We choose the 60-month yield as our long rate because it is available back to 1959:2, whereas longer-term yields such as the 120-month yield become available only in the early 1970s. We use the Federal Funds rate as the short term rate in order to be consistent with the macroeconomic model that we examine in Section 6. The DSGE model does not differentiate between the monetary policy rate and the short-end of the Treasury yield curve (e.g. a 3-month bill rate).\(^9\) To check for robustness, we ran our simulations with alternative

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\(^7\)In addition, one could impose that two shocks, the news shock and innovation to current productivity, account for all of the FEV of TFP. Following BSS, we do not impose this restriction explicitly. Yet, as it turns out, two shocks account for virtually all of the FEV of TFP for all horizons.

\(^8\)At any point in time TFP moves for three possible reasons. First, there may be current innovations to productivity. Second, past news shocks are realized as subsequent movements in productivity. Third, past productivity innovations propagate forward through the lag structure in the VAR.

\(^9\)This approximation seems reasonable since in practice, the Federal Funds rate and short-end bill rates move very closely together. More precisely, the correlation coefficient of the Federal Funds rate and the 3-month bill rate over the 1959:2-2005:2 period is 0.984. The Federal Funds rate is slightly more volatile and has a higher mean than the 3-month bill rate. For our VAR and DSGE exercises, these differences are not important.
measures of the slope and the short rate and found all of the main results to be unchanged.\textsuperscript{10} For the macroeconomic data we use two datasets. The first is a small set of macroeconomic variables consisting of TFP, consumption and inflation. Our measure of TFP is a quarterly version of the series constructed by Basu, Fernald and Kimball (2006). This series exploits first-order conditions from a firm optimization problem to correct for unobserved factor utilization and is thus preferable to a simple Solow residual measure of TFP.\textsuperscript{11} Our measure of consumption is the log of real chain-weighted personal consumption expenditures. For inflation, we use the growth rate of the GDP deflator.

The second dataset is a larger dataset that adds four variables to our smaller dataset. These variables are the log of real chain-weighted GDP, the log of real chain-weighted gross private domestic investment and the log of the S&P 500 composite index deflated by the consumer price index.

All of the macroeconomic series are obtained from the FRED II database of the St. Louis Fed and are available in quarterly frequency. The term structure and stock market data are available in daily and monthly frequency. We convert them to quarterly frequency by computing arithmetic averages over the appropriate time intervals. The sample period is 1959:2-2005:2 (with the start date limited by the availability of 60-month yield). Both the baseline VAR and the extended VAR are estimated in levels with 4 lags of each variable. To improve precision, we impose a Minnesota prior on the estimation and compute error bands by drawing from the posterior.\textsuperscript{12}

\section{What Moves the Slope of the Term Structure?}

In this section we answer the main question of the paper. We do so by first extracting the shocks that explain most of the movements in the slope of the term structure, our ‘target’ variable in the VAR. Second, we look for different possible interpretations of this shock. In

\textsuperscript{10}There are two important alternative measures of the slope. First, we replaced the 60-month yield with the 120-month zero-coupon yield as computed by Gurkaynak, Sack and Wright (2007) and the Federal Funds rate by the 3-month bill rate. Second, we used a Nelson-Siegel style slope factor as computed in Diebold and Li (2006).

\textsuperscript{11}Basu, Fernald and Kimball (2006) also make use of industry level data to correct for differences in returns to scale. Since this industry level data is available only on an annual basis, our quarterly TFP measure does not include this returns to scale correction. See Sims (2009) for details.

\textsuperscript{12}We performed a battery of robustness checks with other macroeconomic variables including data that allowed us to estimate the VAR on monthly frequency. We discuss the responses of some of the added variables in the next section but note that none of the main conclusions is affected by the different changes in VAR specification. Also, we dropped the Minnesota prior and estimated the VAR with OLS instead, computing the error bands by bootstrapping from the estimated VAR. Details are available from the authors upon request.
particular, we pursue the hypothesis that this shock captures news about future innovations to TFP. Third, we show that our results are robust to a variety of alternative assumptions.

4.1 Slope Shocks

As described in Section 2, we extract the shocks that maximizes the fraction of the FEV of the slope that is explained by those shocks. We set the forecast horizon to $0 \leq k \leq 40$ quarters, weighing the importance of each of the forecasts equally. This choice is motivated by the fact that we want to capture short- and medium-run movements in the term structure slope while providing at the same time reliable estimates at the long end of the forecasting horizon. We limit our analysis to two shocks ($n = 2$) because we find that two shocks explain virtually all the movements in the slope. The following results refer to the small VAR described above.

Figure 1 displays the fraction of the FEV of the different variables explained by the first shock. The solid line corresponds to the median estimate, while the dotted lines denote the 16%-84% error bands. As the top left panel shows, this first shock explains more than 85% of all slope movements over the entire 0 to 40 quarter forecast horizon. The second shock (not shown) accounts for virtually all of the remaining fraction of the FEV of the slope. This result is robust across many different VAR specifications. For example, in the extended VAR that we examine at the end of this section, one shock explains about 75% of all slope movements and the second shock accounts for almost all of the remaining 25%. In other words, two shocks are sufficient to understand all movements in the slope and to an approximation, the first shock is by far the most relevant. We thus focus on the properties of this first shock only.\[13\]

The other panels in Figure 1 show that the slope shock also explains about 50% of the Federal Funds rate over the entire horizon, suggesting that slope movements are largely driven by variations in the short end of the term structure. For the macroeconomic variables the slope shock explains very little of variations in TFP, consumption and inflation at short horizons. As the forecast horizon increases, however, the slope shock gradually accounts for a larger fraction of the movements in these variables. In particular, the shock explains more

\[13\] As explained in Section 2, identifying the two most important shocks amounts to finding the innovations associated with the two largest eigenvalues of the matrix $S$ defined in (3); i.e. the first shock corresponds to the eigenvector with the largest eigenvalue and the second shock corresponds to the eigenvector with the second largest eigenvalue. As Uhlig (2003) explains, however, there are two other pairs (or rotations) of eigenvectors that explain together an equal fraction of the total FEV of the slope. We compared all of our results with these two other rotations and found that for each rotation, there exists one shock that explains over 50% of slope movements. This shock has very similar properties than the first shock we consider in the text.
than 40% of the consumption variation at a 20 quarter horizon and about 30% of TFP variations 40 quarters ahead (with this latter fraction increasing towards 50% for forecast horizons beyond 40 quarters). This confirms earlier findings by Ang and Piazzesi (2003), Diebold, Rudebusch and Aruoba (2006) and Evans and Marshall (2007) that there are important linkages between slope movements and macroeconomic fluctuations. Our analysis adds the qualification, however, that these linkages are mostly present for medium- and longer-term macroeconomic fluctuations whereas high-frequency variations in macroeconomic variables are almost completely orthogonal to slope innovations.

The second step in our approach is to provide an economic interpretation of the slope shock. We do this by examining the IRFs of the different variables to an innovation in the slope shock. Figure 2 displays the results. The term spread increases on impact of the shock, while the long end of the term structure remains roughly constant on impact before becoming slightly negative.\textsuperscript{14} The strong reaction of the spread is driven largely by the drop in the Federal Funds rate. Interestingly, the slope shock has no significant impact on either TFP or consumption on impact, but within 2 quarters of the shock, both of these variables start to increase significantly to what appears to be a permanently higher level. Finally, inflation drops significantly on impact of the slope shock and remains below its initial rate for more than 2 years. This drop in inflation is, however, smaller than the drop in the Federal Funds rate, implying that the real Federal Funds rate also turns negative.

How do we interpret this shock? The apparent permanent response of TFP and consumption suggests that the slope shock captures technological innovations leading to an increase in productive capacity in the future. Such a supply-side interpretation also rationalizes why, despite the loosening of monetary policy, inflation falls and remains persistently lower for more than two years. More specifically, the macroeconomic dynamics in Figure 2 look very much like the responses to a news shock about future TFP as identified in Barsky and Sims (2009) and Sims (2009). These two papers report that TFP news shocks lead to a delayed but permanent increase in TFP and consumption and a sharp drop in both inflation and short-term interest rates. Both papers also find that TFP news shocks explain almost none of high-frequency variations in TFP and consumption but account for 40% or more of the two variables at horizons of 20 quarters or more.

Before examining this TFP news interpretation in more detail, it is important to consider whether other prominent macroeconomic shocks are consistent with these IRFs. Monetary shocks are often considered in both macroeconomic studies as well as term structure studies

\textsuperscript{14}The long bond rate is not in our estimated VAR. Its IRF is constructed as the sum of the term spread and the Federal Funds rate.
The monetary shock interpretation appears clearly inconsistent with the IRFs in Figure 2. If the drop in the Federal Funds rate was related to an exogenous monetary policy intervention, then inflation should increase rather than decrease and there should be no permanent effect on either consumption or TFP (e.g. Christiano, Eichenbaum and Evans, 2005). Our technology news hypothesis, by contrast, implies that monetary policy reacts endogenously to the drop in inflation and is thus only indirectly the main driver of the slope.\footnote{This result is consistent with Evans and Marshall (2007) who also conclude that the systematic reaction of monetary policy is an important channel through which macroeconomic shocks affect the term structure.} Taken together, the IRFs to a slope shock are inconsistent with a monetary shock interpretation.

A second type of shock considered in the macroeconomics literature are demand shocks, either in the form of exogenous changes in government deficits (Evans and Marshall, 2007; Dai and Phillippon, 2008) or exogenous changes to the effective interest rate that applies to savings and investment decisions (De Graeve, Emiris and Wouters 2008). Similar to exogenous monetary policy shocks, such demand shocks should not have a permanent positive effect on either consumption or TFP. Likewise, we know of no theory of demand shocks that produces a prolonged decline in both inflation and the Federal Funds rate in response to a positive demand shock.

A third type of shock from the macro-labor literature is a shock to the marginal rate of substitution (MRS) between consumption and leisure. Evans and Marshall (2007) study the impact of this shock on the term structure and find that this shock has a statistically insignificant affect on the slope and inflation while increasing both real activity and the Federal Funds rate. These predictions are inconsistent with the IRFs in Figure 2. We conclude that MRS shocks cannot be an interpretation of our slope shock.\footnote{Evans and Marshall (2007) find that MRS shocks are primarily important for variations in the level of the term structure but have no significant impact on the slope.}

A fourth type of macroeconomic shocks is a contemporaneous innovation to TFP as traditionally assumed in the business cycle literature. We identify a contemporaneous TFP shock by ordering TFP first in our VAR and extracting the first column of a Cholesky decomposition. Figure 3 display IRFs to this contemporaneous TFP shock. Notably, TFP rises dramatically on impact and so does consumption. Both of these responses are persistent but ultimately transitory. Furthermore, the shock has no significant effect on the term spread and only a delayed but negligible effect on the Federal Funds rate. All of these IRFs are inconsistent with our slope shock, thus suggesting that it is indeed news about future productivity innovations that are a main driver of the slope of the term structure.
4.2 Slope Shocks are News Shocks

We now refine the TFP news interpretation of the slope shock by formally identifying a news shock. News shocks about future productivity have recently been resuscitated as a potential source of business cycle fluctuations in recent work by Beaudry and Portier (2006), Jaimovich and Rebelo (2009) and Schmidt-Grohe and Uribe (2008), among others. A news shock is information about the future level of TFP. Beaudry and Portier (2006) model the process for TFP as:

\[ a_t = v_t + D_t, \]  

where \( a_t \) is the log of TFP; and \( v_t \) and \( D_t \) are two independent exogenous components. The component \( v_t \) captures potentially persistent but transitory surprise movements in TFP. The component \( D_t \) is non-stationary and assumed to follow a distributed lag process in past innovations; i.e. \( D_t = d(L)\eta_t \) with \( d(0) = 0 \) and \( d(1) = 1 \). Innovations \( \eta_t \) are interpreted as news shocks about future productivity because they do not affect TFP contemporaneously, but only with a delay of one or more periods.\(^{17}\)

Rather than following the empirical approach of Beaudry and Portier (2006) who identify news shock with a mix of short- and long-run restrictions on stock prices and TFP, we adopt the more recent identification approach proposed by BSS. As described in the previous section, the BSS approach is similar in spirit to our statistical extraction of the slope shock and consists of identifying the shock that explains most of TFP variations over a given forecast horizon but is orthogonal to contemporaneous innovations in TFP. As such, the BSS identification satisfies the definition of news about future productivity (i.e. \( d(0) = 0 \)) in Beaudry and Portier (2006) and also allows for news to have a permanent effect on TFP (i.e. \( d(1) = 1 \) is possible but not required).

Figure 4 displays the fraction of the FEV of the variables in the VAR explained by the TFP news shock. As we found for the slope shock, the TFP news shock explains almost none of the movements in macroeconomic variables on impact but up to 50% of consumption variations after 20 quarters and about 40% of TFP variations after 40 quarters. The shock also explains over 60% of term spread movements over all horizons and between 60% and 80% of Federal Funds rate movements. In other words, the TFP news shock seems to be a major driver of term structure movements.\(^{18}\)

\(^{17}\)TFP news shocks resemble recent characterizations of technological adoption by Rotemberg (2003) or Comin and Gertler (2006). While neither of these papers imposes the zero restriction on impact of the shock, they both argue that it takes on average several years for new technologies to be adopted even though these innovations are known to exist and be commercially valuable. See Rotemberg (2003) for an extensive discussion of evidence about the slow diffusion of technological innovations.

\(^{18}\)Figure 4 also shows that the TFP news shock explains almost nothing of high-frequency fluctuations
Figure 5 reports the IRFs of the different variables to the TFP news shock (solid blue lines) and reproduces the IRFs to the slope shock from Figure 2 for comparison (dashed red lines). The similarity in results is striking. In particular, the slope jumps up significantly on impact and then returns back to its pre-shock value after 10 to 15 quarters; TFP increases gradually from zero (by construction for the news shock identification) to a permanently higher level (even though no constraint on long-run effects is imposed); consumption increases slightly (but insignificantly) on impact and then gradually increases to a permanently higher level; and both inflation and the Federal Funds rate drop markedly on impact and remain below their initial value for more than 15 quarters. As with the slope shock, the drop in the Federal Funds rate is larger than the drop in inflation, implying a decline in the real Federal Funds rate.

To further illustrate the correspondence between the TFP news shock and our slope shock, we extract the time series of each of the two shocks and plot them together. As Figure 6 shows, the slope shock is slightly more volatile than the TFP news shock but overall, the two shocks move closely together. In fact, the correlation of the two is 0.87. This close correspondence is surprising because the identification criteria behind the two shocks are very different from each other. The slope shock is extracted by maximizing the FEV of the slope while the TFP news shock is extracted by maximizing the FEV of TFP subject to the additional constraint that the shock is orthogonal to contemporaneous TFP movements. Hence, there would be no a priori reason to believe that the two innovations capture the same economic shock.

Finally, to assess the empirical relevance of the TFP news shock, Figure 7 plots the historical time series of the different variables in the VAR against the simulated times series conditional on the TFP news shock (i.e. assuming that TFP news shocks are the only stochastic innovation). As the two top panels show, TFP news shocks explain very little of the high-frequency fluctuations in TFP and consumption, which is consistent with our conclusions from the FEV decompositions. TFP news shocks also miss most of the high-frequency variations in inflation but capture quite a lot of the medium-frequency movements in inflation, especially during the 1970s and early 1980s. TFP news shocks do a surprisingly good job in accounting for fluctuations in the Federal Funds rate and the slope. In particular, TFP news shocks account for almost all of the large swings in the slope during the 1970s.

Barsky and Sims (2009), by contrast, report that TFP news shocks explain more than 60% of high-frequency variations in inflation and between 40-55% at horizons of 4 quarters and higher. This difference is due to the fact that they compute inflation from the CPI deflator while we use the GDP deflator. We prefer the latter because it represents a broader measure of aggregate prices, does not suffer from substitution bias, and is less affected by large temporary swings in food and energy prices.
and also rationalize the increase in term spreads during the early 1990s and the early 2000s. This close fit is striking and leads to two important lessons. First, a large part of Federal Funds rate fluctuations are driven by news about future supply-side innovations. Second, through the endogenous response of the Federal Funds rate, TFP news shocks are a main driver of the slope of the term structure.

4.3 Robustness

In this section we show that our empirical results are robust to a number of possible concerns. The first potential issue with our results concerns mismeasurement of technological progress. In particular, advances in technology may not come through increases in TFP but rather through technological progress that is embodied in new capital. Hence, if capital services are not appropriately measured, our identification may mistake embodied (i.e. capital-specific) technological progress for TFP improvements. This concern is motivated by recent empirical evidence from Fischer (2005) who reports that embodied technological shocks are a main driver of business cycle fluctuations. To address this issue, we add Fischer’s (2005) relative price deflator series for investment and equipment goods to our VAR and rerun both the slope shock identification and the TFP news shock identification. In response to the slope shock, both relative price deflators increase slightly on impact and then decrease significantly after about 10 quarters to a permanently lower level. In response to the TFP news shock, by contrast, neither of the relative price deflators reacts significantly. All of the other results remain unaffected. This suggests, on the one hand, that TFP news shocks are not erroneously capturing capital-specific embodied technological progress. On the other hand, the slope shock seems to picks up not only news about future TFP increases, but also news about future embodied technological progress. This could be one of the reasons why the extracted slope shock is slightly more volatile than the TFP news shock.

A second issue is the extent to which our results are robust to alternative VAR specifications. We estimated many different VAR specifications and found our results to be generally robust. For space reasons, we report here only one of these alternative specifications, which extends the baseline VAR with output, investment and the S&P 500 composite index. We choose this particular extension because it allows us draw comparisons with the recent empirical literature on news shocks and because it provides a useful benchmark for the DSGE model that we introduce in Section 6. Figure 8 reports the IRFs to the TFP news shock for this extended VAR. As in the smaller VAR, the TFP news shock has a gradual but

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19The relative price series we use are updated by DiCecio (2008).
20In the interest of conciseness, we do not plot the fraction of FEVs explained by TFP news shocks for
permanent effect on real variables. Consumption now increases somewhat on impact of the shock. Output declines slightly on impact, but the change is not very significant. Investment, by contrast, contracts significantly over the first two periods. The real stock market index increases on impact and remains significantly higher for about four years before slowly returning back to its initial value. Finally, both inflation and the Federal Funds rate drop markedly on impact and remain persistently below their initial value for 15 to 20 quarters. Since the long rate barely moves, the term spread increases on impact of the shock and then gradually returns to its average value. Overall, these results look very similar to the results obtained above with the baseline VAR.

The small inverse reaction of consumption and output on impact of the TFP news shock matches closely the findings in Sims (2009). At the same time, these results contradict Beaudry and Portier (2006) who find that consumption and real activity (measured by either hours or investment) both display large positive reactions almost immediately after a TFP news shock. Furthermore, Beaudry and Portier’s (2006) TFP news shocks account for a large part of the high-frequency fluctuations in real aggregates whereas this is not the case in our analysis. As Sims (2009) shows, this difference in results is due to the different identification approach employed by Beaudry and Portier (2006). They identify a TFP news shock either as the VAR innovation that may have a permanent long-run effect on TFP, or the innovation that is orthogonal to current TFP but may affect stock prices contemporaneously.

There are several advantages to the BSS identification approach over the ones employed by Beaudry and Portier (2006). First, long-run restrictions have been shown to have very poor finite sample properties (e.g. Faust and Leeper, 1997). According to Monte-Carlo simulations by Francis, Owyang and Roush (2007), FEV-based criteria such as the one employed by BSS perform significantly better at identifying technology shocks. Second, the BSS approach imposes that news shocks account for the maximum variation in TFP over the entire short- to medium-run horizon. The BSS approach is thus more inclusive than a long-run restriction and directly addresses the problem that shocks identified by long-run restrictions commonly account for only a modest fraction of TFP fluctuations even at forecast horizons of 10 or more years. Third, stock prices react to many different shocks and the different variables of this extended VAR. Interestingly, the TFP news shock accounts for an even larger fraction of TFP and consumption movements at the 20-40 quarter horizon. Similarly, the shock explains almost nothing of output and investment fluctuations on impact but about 50% of both variables after 20 quarters and more. For the term structure, in turn, the shock explains between 40% and 50% of movements in the slope and the Federal Funds rate over the entire horizon. This is somewhat less than in the baseline VAR but still very sizable. Finally, the TFP news shock explains roughly 20% of inflation and stock market movements over the entire horizon.

21Sims (2009) also reports that hours worked decline for the first few quarters after the TFP news shock. We find the same result if we include hours worked as an additional variable in the VAR.
are thus a relatively uninformative measure of future technology innovations. For all these reasons, we prefer the BSS approach to identify TFP news shocks.

5 Term Premia versus the Expectations Hypothesis

Our VAR framework allows us to decompose the reaction of the long rate into variations due to term premia and expectations about future short rates (i.e. the Expectations Hypothesis). We can decompose the yield on a $T$-period yield $r^T_t$ (in our case the 60-month yield) as

$$ r^T_t = \frac{1}{T} \sum_{i=0}^{T-1} E_t r_{t+i} + t_p t, \quad (6) $$

where the $E_t r_{t+i}$ are time $t$ expectations of future short rates; and $t_p t$ denotes term premia. The reaction of the long-rate with respect to TFP news shocks may be relatively small either because the Expectations Hypothesis part $1/T \sum_{i=0}^{T-1} E_t r_{t+i}$ and the term premia part do not respond strongly or because variations in the two almost cancel each other out. This, in turn, determines the importance of term premia fluctuations for the reaction of our slope measure.

The technical difficulty with this decomposition is that term premia are inherently unobservable. Here, we follow Campbell and Shiller (1987, 1991) and use our larger VAR to compute expectations of short rates conditional on TFP news shocks. The term premia response to the TFP news shock is then simply the difference between the actual long rate response to the shock and the response as implied by the Expectations Hypothesis computed from the VAR. Figure 9 shows the resulting decomposition both for the long rate (top panels) and the spread (bottom panels). As the top panels show, term premia react positively to a TFP news shock before turning slightly negative after about 10 quarters. This initial jump in term premia occurs because the long rate under the Expectations Hypothesis displays a more marked drop on impact than the actual long rate. Or put differently, the reaction of the long rate to the TFP news shock is relatively small because term premia variations neutralize almost all of the initial drop in the long rate as implied by expectations of future short rates. As the bottom panels show, this implies that almost half of the initial jump in the slope is due to the increase in term premia. The response of the slope under the Expectations Hypothesis (i.e. as implied by short-rate fluctuations) remains, however, significant and returns only gradually back to its initial value. Hence, the endogenous reaction of the Federal Funds rate remains a quantitatively important direct channel through which TFP news shocks affect the slope.

The large and significant reaction of term premia is consistent with the general statistical
rejection of the Expectations Hypothesis in the finance literature (e.g. Fama and Bliss, 1987; Campbell and Shiller, 1991; Cochrane and Piazzesi, 2005). At the same time, the Expectations Hypothesis by itself can account for more than half of the slope response to a news shock and thus remains empirically relevant, which is consistent with the basic message of Campbell and Shiller (1987) and more recently King and Kurmann (2002).

6 A DSGE-News Model of the Term Structure

Our final exercise is to evaluate how well a linearized New Keynesian DSGE model along the lines of Smets and Wouters (2007) can account for term structure movements in response to TFP news shocks. The linear approximation makes it easy to solve and estimate the model despite its relative complexity. Unfortunately, the linearization coupled with the assumption of homoscedastic innovations implies constant term premia. Since time varying term premia are important for understanding long bond yield movements we combine the linearized DSGE model with an affine formulation of the pricing kernel that allows for time-varying risk. Under no arbitrage, long bond yields can then be derived recursively as a combination of expected future short rates and time-variation in term premia. Movements in both components are governed entirely by the dynamics of the states variables from our DSGE model, which imposes considerable discipline on their dynamics.

6.1 Model

The macro part of the model is very similar to the one presented in Smets and Wouters (2007) and contains several real and nominal frictions. Specifically, the model features sticky nominal price and wage setting that allows for indexation to lagged inflation, external habit persistence in consumption, investment adjustment costs, variable capital utilization and fixed costs of production.

The main structural difference of our model to the one presented in Smets and Wouters (2007) is that we specify TFP as an exogenous process with a stochastic trend, driven by both a contemporaneous shock and a news shock; i.e.

$$\mu_t = (1 - \gamma)\mu + \gamma\mu_{t-1} + \varepsilon^{current}_t + \varepsilon^{news}_{t-j},$$

where $\mu_t = a_t - a_{t-1}$ is the growth rate of TFP; $\varepsilon^{current}_t$ is the contemporaneous shock; and $\varepsilon^{news}_{t-j}$ is the news shock. This news shock impacts actual TFP in period $t$ but is known $j$ periods in advance. Both the contemporaneous shock and the news shock are i.i.d. processes.
with mean zero and variance $\sigma_{\text{current}}^2$ and $\sigma_{\text{news}}^2$. In our VAR, TFP begins to react after the first period after the news shock. So we set $j = 1$. We also use a slightly different description of monetary policy than Smets and Wouters (2007) and specify the short-term nominal rate $r_t$ as a function of expected inflation $E_t\pi_{t+1}$, the output gap $y_{\text{gap},t}$ and output growth $\Delta y_t$ (rather than growth in the output gap)

$$r_t = \rho r_{t-1} + (1 - \rho)[r + \theta\pi E_t(\pi_{t+1} - \pi) + \theta y_{\text{gap},t} y_{\text{gap},t} + \theta \Delta y(\Delta y_t - \Delta y)],$$

(8)

where the output gap is defined as the difference between actual output and potential output if there were no nominal price and wage rigidities.

For space reasons, we relegate a detailed description of the full model and its linearization to the appendix. The Rational Expectations equilibrium of the resulting system of equations is computed using the numerical algorithm of King and Watson (1998) and can be expressed as a linear state-space system

$$\mathbf{Y}_t = \phi_\mathbf{Y} + \Phi_\mathbf{Y} \mathbf{S}_t$$

(9)

$$\mathbf{S}_t = \phi_\mathbf{S} + \Phi_\mathbf{S} \mathbf{S}_{t-1} + G \varepsilon_t.$$  

(10)

where the $n \times 1$ vector $\mathbf{Y}_t$ contains the endogenous variables; the $k \times 1$ vector $\mathbf{S}_t$ contains the states; and the $k_x \times 1$ vector $\varepsilon_t$ contains the i.i.d. innovations to the exogenous shocks that we assume multivariate normal $(0, I)$.\(^{22}\)

For the term structure part of the model, notice that the short-term yield is part of the macro system and therefore included in the linear state-space system; i.e.

$$r_t = \delta_0 + \delta_1' \mathbf{S}_t,$$

(11)

where $\delta_0$ and $\delta_1'$ contain the appropriate elements of $\phi_\mathbf{Y}$ and $\Phi_\mathbf{Y}$, respectively. The yield on a $T$-period discount bond is defined as

$$r^T_t = -\frac{\log P^T_t}{T},$$

(12)

where $P^T_t$ is the period-$t$ price of the bond with $P^0_t = 1$. Under no arbitrage, this price satisfies

$$E_t[M_{t+1}^S P^T_{t+1}^{-1}] = P^T_t.$$  

(13)

\(^{22}\)There are $k_y = k - k_x$ endogenous states (i.e. predetermined endogenous variables), which are ordered first in $\mathbf{S}_t$. Hence, $G$ is a $k \times k_x$ matrix with zeros in the upper $k_y \times k_x$ block and a diagonal matrix with the exogenous shocks’ standard deviations in the lower $k_x \times k_x$ block.
where $M^8_{t+1}$ is the nominal pricing kernel. Following Ang and Piazzesi (2003) and many others in the latent factor no-arbitrage literature, we assume that the logarithm of this pricing kernel is described by

$$
\log M^8_{t+1} = -r_t - \frac{1}{2} \Lambda_t' \Lambda_t - \Lambda_t' \varepsilon_{t+1},
$$

(14)

where the $k_x \times 1$ vector $\Lambda_t$ denotes the market price of risk associated with the different shocks in $\varepsilon_t$. These risk factor are assumed to follow an affine process in the states

$$
\Lambda_t = \lambda_0 + \lambda_1 S_t,
$$

(15)

with the $k_x \times 1$ vector $\lambda_0$ defining average risk; and the $k_x \times k$ matrix $\lambda_1$ defining how risk varies depending on the state of the economy. Given (9)-(15), bond prices can be computed recursively as linear functions of $S_t$ that can be decomposed into fluctuations due to expected future short rates (i.e. the Expectations Hypothesis) and time variations in term premia (see the appendix for details).

As shown by Wu (1996) and Bekaert, Cho and Moreno (2010), the pricing kernel implied by linearized DSGE models with homoscedastic innovations represent a special case of the formulation in (14) with $\lambda_0$ a function of the different structural parameters of the DSGE model and $\lambda_1 = 0$. In other words, the linearized DSGE model implies that risk and therefore term premia are constant. To allow for time-variation in term-premia we let $\lambda_1$ be non-zero. This potentially involves estimating the entire $k_x \times k$ matrix $\lambda_1$, which is large for our DSGE model. To make the estimation manageable, we impose two restrictions. The first restriction is that we let risk vary only with respect to two macro variables: expected inflation $E_t \pi_{t+1}$ and expected changes in marginal utility of consumption $E_t \Delta u_{c,t+1}$. This allows us to express risk associated with the news shock as

$$
\lambda^\text{news}_t = \lambda^{\text{news}}_{0} + \tilde{\lambda}^{\text{news}}_{1,\pi} E_t \pi_{t+1} + \tilde{\lambda}^{\text{news}}_{1,\Delta u_c} E_t \Delta u_{c,t+1},
$$

(16)

where $\tilde{\lambda}^{\text{news}}_{1,\pi}$ and $\tilde{\lambda}^{\text{news}}_{1,\Delta u_c}$ tell us how the price of risk with respect to the news shock reacts to changes in expected inflation and changes in the expected change of the marginal utility of consumption, respectively. Both of these variables are part of the state-space solution of our DSGE model in (9)-(10). The second restriction follows naturally from our limited information estimator in that the only exogenous shock we consider is the news shock. As the appendix shows in detail, the two restrictions together imply that $\tilde{\lambda}^{\text{news}}_{1,\pi}$ and $\tilde{\lambda}^{\text{news}}_{1,\Delta u_c}$ are the only additional free parameters to estimate. This imposes considerable discipline on our
estimation (by comparison Ang and Piazzesi, 2003 estimate a total of 13 different coefficients in their formulation of $\Lambda_t$).

There are two important features of our modelling of risk. First, our formulation of time-varying risk can be motivated by the consumption-based asset pricing literature, which says that changes in risk over the business cycle must come from changes in the conditional covariances between inflation and the marginal utility of consumption (e.g. Piazzesi and Schneider, 2006). Establishing the link between risk and conditional covariances explicitly in the context of a non-linear DSGE model has been largely unsuccessful. Our formulation should therefore be considered as a basic test of whether variations in risk as a linear function of two macro variables are capable of generating quantitatively large term premia fluctuations. Second, the macro dynamics of our model as described by the state-space system in (9)-(10) are independent of time-variation in risk.23 Since risk depends on the macro states, however, the joint estimation of both macro and term structure dynamics imposes discipline on the parameters of the macro model.

6.2 Estimation

We partition the parameters of the model into two groups. The first group consists of parameters that we calibrate to match long-run moments of the data. The second group is estimated to match the IRFs to a news shock as generated by the empirical VAR. All values reported are for a quarterly frequency.

Table 1 presents the calibrated parameters. The growth rate of output $\Delta y$ and TFP $\mu$ are set to match the average growth rate of real GDP and TFP in the data (1.86% and 1.29% annually for the 1959-2005 sample). The next four parameters imply a labor share of 0.675 in line with Gollin (2002), an average annualized quarterly real interest rate of 2.34% as measured in our data; an annual depreciation rate of 10 percent; and an average markup for final goods producers of 11% as reported by Basu and Fernald (1997). The elasticity of substitution across labor $\theta_w$ is set as in Smets and Wouters (2007); the unit elasticity of labor supply $\eta$ is a compromise between values suggested in the microeconomic and macroeconomic literatures; and the fixed cost in production (not reported here) is set so that economy-wide net profits are zero as suggested by Basu and Fernald (1994) or Rotemberg and Woodford (1995).

The second group of parameters is estimated by minimizing a weighted distance between the model-implied IRFs to a news shock and the empirical counterparts from the VAR.

23This is consistent with the modelling of time-varying shocks in Justiano and Primiceri (2008) where changes in variance per se have no impact on the decision rules in the model.
Specifically, denote by $\hat{\Psi}$ a vector of empirical IRFs to a news shock over obtained from a VAR. Likewise, denote by $\Psi(\zeta)$ the same vector of IRFs implied by the model, where $\zeta$ contains all the structural parameters of the model. The estimator for the second group of parameters $\zeta_2 \subseteq \zeta$ is

$$\hat{\zeta}_2 = \arg \min_{\zeta_2} \left[ \hat{\Psi} - \Psi(\zeta) \right]' \Omega^{-1} \left[ \hat{\Psi} - \Psi(\zeta) \right],$$

where $\Omega$ is a diagonal matrix with the sample variances of $\hat{\Psi}$ along the diagonal. This limited-information approach is the same than the one implemented by Christiano, Eichenbaum and Evans (2005) for a monetary policy shock. Here, we adapt it for our purposes by first estimating the parameters governing the response of TFP to an exogenous news shock and then, in a second step, by estimating the remaining structural model parameters such as to match the IRFs of other variables in the VAR. We adopt this two step approach because we want to evaluate the ability of our model to generate realistic term structure and macroeconomic dynamics to a news shock given the evolution of observed TFP.

On the macro side, we include the IRFs of TFP and all four macroeconomic variables of our extended VAR in the objective function. On the term structure side, we include the IRFs of the short rate, the long bond rate and the term spread as well as the IRFs of the long bond rate and the term spread implied by the Expectations Hypothesis (as computed in the previous section from the VAR). This provides us with a total of 10 empirical IRFs. For each of these IRFs, we include the entire 40 quarter horizon in the estimation criteria.

### 6.3 Results

We provide two estimates of the parameters of the model. The first is an unconstrained estimation where we let the estimated parameters in $\zeta_2$ take on any value within the theoretically admissible bounds. The second is a constrained estimation where we force a subset of parameters in $\zeta_2$ to not exceed values commonly found in the DSGE literature. The first set of estimates is motivated by our desire to maximize the fit of the model. The second set is suggestive of the difficulties of matching the responses of a macro and finance variables to a news shock while maintaining an ability to match other known empirical facts.

Figure 10 plots the model IRFs implied by the unconstrained estimation and compares them to the IRFs from the VAR (with the grey-shaded areas demarking the 16%-84% error bands of the VAR responses). Overall, the estimated model does well in matching the responses of the different macro variables. As the plot for TFP shows, the stochastic growth process in (7) almost perfectly traces the gradual increase of TFP after the news shock.
The model also matches closely the initial jump in consumption and the subsequent gradual increase to the new balanced growth level. For output and investment, the model misses the initial drop in both variables and fails to generate the sharp increase in investment over the following periods. Over the longer run, however, the model matches the increase of both variables to the new balanced growth level. Likewise, the model generates the overall shape of the inflation response even though on impact of the news shock, the model implies a drop in inflation that is too modest.

For the term structure variables, the performance of our model is more mixed. The model is successful in generating the sizable jump in the spread and the gradual return back to its average value that we obtain from the VAR. In order to get this large initial jump, however, the model needs an increase in the term premium that is too large relative to its empirical counterpart. As a result, the long-bond yield initially responds with the wrong sign before matching the data. For the short rate, the model generates a sharp drop on impact of the shock but this drop is only about half as large as in the VAR. Consequently, the initial response of the spread implied by the Expectations Hypothesis remains well below its VAR counterpart.

The estimated parameters that generate the IRFs in Figure 10 are reported in Table 2 in the column labeled 'unconstrained estimates'. The estimates $\kappa_p = 1$ and $\omega_p = 0$ indicate that the data favors a purely forward-looking New Keynesian Phillips curve (NKPC) with little price rigidity (i.e. with standard Dixit-Stiglitz goods differentiation, a coefficient of $\kappa_p = 1$ implies an average price duration of only 1.6 quarters). By contrast, the data requires an extreme degree of nominal wage rigidity with an estimated frequency of wage reoptimization of only $1 - \xi_w = 0.01$ per quarter and a degree of indexation for non-reoptimized wages to past inflation of $\omega_w = 0.70$. The main force behind these estimates is the sharp drop of inflation on impact of the news shock, which the model can generate only if inflation is a mainly forward-looking process that reacts strongly to current and future expected marginal

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24 As we mentioned in Section 4, there is controversy about whether output and investment decreases on impact of a news shock. While important, this question is not the focus of our investigation. We therefore attach relatively little importance to whether our model is capable of generating a negative initial reaction of output and investment or not.

25 Since we use a limited information approach to estimate a relatively large number of parameters, it is not surprising that we face a number of weak identification issues. For the unconstrained estimation, this manifests itself in a marginal cost coefficient $\kappa_p$ that tends to wander off towards very large values without much improvement in the estimation objective and, consequently, very large associated standard errors. We therefore fix $\kappa_p = 1$ in this estimation. This does not change any of the conclusions. Also note that the inflation indexation parameter $\omega_p$ is estimated at its lower bound. Since it would not be meaningful to report a standard error at this boundary, we fix the parameter when computing standard errors for the other estimates. We adopt the same approach for any other parameter that is estimated at its respective lower or upper bound.
cost (i.e. $\omega_p$ is small and $\kappa_p$ is large). Marginal cost, in turn, depends positively on wages and negatively on TFP. After a news shock, the negative income effect on labor supply from consumption smoothing puts upward pressure on wages and thus on marginal cost. In subsequent periods, as the expected increase in TFP realizes, marginal cost falls. The drop in inflation on impact and the gradual response thereafter occurs only if there is a lot of wage rigidity (i.e. $\xi_w$ and $\omega_w$ large) so that the initial increase in marginal cost is relatively modest and its negative reaction after the TFP shock realizes is large.

The estimation also has strong implications for capital utilization and investment adjustment cost. The parameter governing the variability of capital utilization $\sigma_u$ is estimated close to its lower bound of 0, which implies that capital utilization is roughly proportional to the rental rate of capital.\(^{26}\) As Dotsey and King (2006) show, variable capital utilization reduces the sensitivity of marginal cost. Hence, the smaller the cost of utilization, the less pressure production exerts on marginal cost. This helps the model reconcile the large expansion of production with the persistent drop in inflation in the wake of the news shock. The investment adjustment cost parameter, in turn, is estimated at its lower bound of $S'' = 0$ (i.e. adjustment costs are zero in the vicinity of the steady state). This estimate is driven by the initial drop of investment and the need for little pressure on marginal cost on impact of the shock. If investment adjustment costs were large, then there would be a strong incentive to smooth investment, which in turn would put upward pressure on production and inflation.

Turning to monetary policy, the estimates $\rho = 0.0$ and $\theta_\pi = 2.92$ indicate that the Fed does not smooth its policy rate and reacts aggressively to inflation. Both parameter estimates are crucial to generate the sharp drop in the Federal Funds rate on impact.\(^{27}\) The estimates $\theta_{ygap} = -0.02$ and $\theta_{\Delta y} = 2.13$ imply that the Fed does not respond to the output gap but reacts strongly to output growth, which is consistent with the findings in Orphanides (2005). The focus on output growth rather than the output gap turns out to be crucial for the model to generate a fall in the Federal Funds rate. In response to a news shock, the output gap in the model increases whereas output growth falls.\(^{28}\) Hence, if monetary policy responded

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\(^{26}\)For $\sigma_u = 0$, depreciation increases linearly with utilization. In the absence of investment adjustment cost (i.e. $S'' = 0$), $\sigma_u = 0$ is inconsistent with the stationarity assumption for interest rates. See the appendix for details. We therefore impose a lower bound of $\sigma_u = 0.001$ on the estimation.

\(^{27}\)The estimate of $\rho = 0$ is not necessarily inconsistent with the literature. For example, Rudebusch (2006) argues that the persistence of the Federal Funds rate in response to a monetary policy shock is better explained by persistence in the exogenous shock process than persistence in the policy rule itself.

\(^{28}\)As described above, the output gap is defined as the difference between actual output and potential output in the absence of nominal price and wage rigidities. In response to a news shock, prices drop abruptly, which means that firms’ average markups decrease. Hence, actual output drops less than potential output (for which markups are constant by definition) and the output gap jumps up. Simulations with more basic policy rules that feature only the output gap show that the fit of the model under this specification falls apart almost completely, with the Federal Funds rate and the slope hardly responding to the news shock.
strongly to the output gap, this would reduce (or even reverse) the already insufficient drop in the Federal Funds rate. A strong response to output growth, by contrast, reinforces the accommodative stance of the Fed thus bringing the model closer to the observed term structure dynamics.29

Finally, consider the estimates of the risk loadings on expected inflation $\hat{\lambda}_{1,\pi}^{\text{news}}$ and the expected change in marginal utility from consumption $\hat{\lambda}_{1,\Delta u_c}^{\text{news}}$. Both coefficients are negative and highly significant. This provides another piece of evidence against a pure form of the Expectations Hypothesis and suggests that a linearized DSGE model on its own is incapable of generating sufficiently large term structure movements.

The estimates for the marginal cost coefficient of $\kappa_p = 1$ and the degree of wage rigidity $\xi_w = 0.99$ exceed other limited- and full-information estimates for the two coefficients in the literature, which typically reports values for $\kappa_p$ around 0.025 and for $\zeta_w$ around 0.75.30 To assess the model’s performance conditional on more realistic price and wage adjustment processes, we fix $\kappa_p = 0.05$ and $\zeta_w = 0.85$. Both of these values are upper bounds found in the literature. The other parameters of the model are reestimated and are reported in the column labeled ‘constrained estimates’ of Table 2. The NKPC is still estimated to be completely forward-looking (i.e. $\omega_p = 0$) and the degree of wage indexation to past inflation goes to its upper bound of $\omega_w = 1$. Also, the model still favors highly variable capital utilization and no adjustment cost to investment. The intuition for these estimates is as above: for inflation to fall on impact of the shock, the NKPC needs to be driven by current and future marginal cost. Marginal cost, in turn, needs to be insensitive to the initial upward pressure coming from the negative income effect on labor supply. For the interest rate rule, there are relatively important changes, indicating that there is an important interplay between price and wage rigidity and monetary policy. In particular, the interest rate rule now exhibits substantial persistence (i.e. $\rho = 0.71$) and the response to output growth is about four times smaller.31

Figure 11 plots the IRFs for the reestimated model. The model still generates an initial jump in consumption but can no longer match the subsequent increase to the new balanced growth level. By contrast, the model now implies a marked drop in output and investment

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29 Barsky and Sims (2009) argue in favor of a similar monetary policy rule that does not respond to the output gap. However, their argument is somewhat different, based on their empirical result that real short-term rates in response to a TFP news shock are positive. As we pointed out above, however, real short-term rates are negative after a TFP news shock if inflation is measured by the more inclusive GDP deflator rather than the CPI deflator.

30 For example, Smets and Wouters’ (2007) full-information estimation yields $\kappa_p = 0.021$ and $\zeta_w = 0.73$.

31 In this ‘constrained estimation’, the issue of weak identification mentioned above manifests itself in an interest rate rule that has a tendency for $\theta_\pi$ and $\theta_{\Delta y}$ to wander off to unreasonably large values. We therefore fix $\theta_\pi = 3$. None of the results change when we fix $\theta_\pi$ to other values above 3.
on impact of the shock. The model also generates inflation dynamics that are reasonably close to the VAR counterpart. On the term structure side, however, the model now performs markedly worse. While the short rate still drops on impact, this drop is only about a third of what we see in the VAR. As a result, the spread implied by the Expectations Hypothesis barely increases and remains well below the VAR response. The model thus requires an even larger term premium on impact of the shock to generate a sufficiently large increase in the observed spread, which means that there is an even larger overshooting of the long rate on impact.

In sum, once we restrict the model to realistic parameter values for wage and price adjustment dynamics, the model has considerable trouble matching the quantitative responses of term structure variables to a news shock. Modern New Keynesian DSGE models as proposed by Smets and Wouters (2007) thus fail to match the quantitative response of both macro and term structure variables to a TFP news shock. Since this class of DSGE models is commonly used for monetary policy analysis this failure represents an important challenge for modern macroeconomics.

7 Conclusion

In this paper we provide a new structural interpretation of the relationship between the slope of the term structure and macroeconomic fundamentals. Our results show that there exists one single shock that can account for a large part of slope movements at all horizons. We interpret the slope shock as a news shock about future innovations to TFP. We assess this interpretation formally using the identification of Barsky and Sims (2009) and Sims (2009) and find a striking correspondence. In response to a positive news shock real activity does not initially respond. At longer horizons real activity increases gradually towards a higher permanent level. Inflation falls sharply on impact of the shock and returns only slowly to its initial level. Monetary policy reacts to the low inflation by lowering the Federal Funds rate. As a result, the short-end of the term structure falls, which in turn leads to a sharp increase in the slope. Endogenous monetary policy thus provides an important channel through which TFP news shock transmit to movements in the slope.

Our news shock provides a structural interpretation for why the yield curve is such a reliable predictor of future output growth. These movements in the slope are asset markets responding to news about the future level of productivity. Future productivity is of course a

\[32\] It is possible to find parameter combinations that allow the model to match the responses of the term structure data almost perfectly. The problem is that for these parameter combinations the model’s performance for the macroeconomic variables deteriorates significantly.
main determinant of future output. This result also provides a structural interpretation for
the result in the existing macro-finance literature on the strong linkages between the yield
curve, inflation, and real output.

Our results provide an important benchmark to evaluate theories of the term structure
and, more generally, DSGE models. We show that a medium-scale DSGE model along the
lines of Smets and Wouters (2007) falls well short of matching the term structure response to
a TFP news shock that we see in the data. This failure of generating realistic term structure
dynamics is problematic for two reasons. First, asset prices (of longer-term securities in
particular) are an important determinant of consumption and investment decisions. If a
DSGE model cannot simultaneously match both macroeconomic and asset price dynamics,
then this suggests a serious empirical shortcoming of theory. Second, medium-scale DSGE
models are increasingly used for monetary policy analysis. If these models fail to generate
reasonable term structure dynamics, then it seems difficult to trust them for the evaluation
of how monetary policy transmits into the economy. A fruitful path for future research is
to search for a mechanism to augment DSGE models that can generate large endogenous
variations in term premia and to estimate these models in a full-information context with
both term-structure data and news shocks.
References


Table 1: Calibrated parameters

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Description</th>
<th>Calibration</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta y$</td>
<td>Steady state growth rate of output</td>
<td>1.0032</td>
</tr>
<tr>
<td>$\mu$</td>
<td>Steady state growth rate of TFP</td>
<td>1.0046</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>Elasticity of production to labor</td>
<td>0.75</td>
</tr>
<tr>
<td>$\beta$</td>
<td>Discount factor</td>
<td>0.997</td>
</tr>
<tr>
<td>$\delta$</td>
<td>Depreciation rate</td>
<td>0.025</td>
</tr>
<tr>
<td>$\theta_p$</td>
<td>Elasticity of substitution across goods</td>
<td>10</td>
</tr>
<tr>
<td>$\theta_w$</td>
<td>Elasticity of substitution across labor</td>
<td>3</td>
</tr>
<tr>
<td>$\eta$</td>
<td>Frisch elasticity of labor supply</td>
<td>1</td>
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Table 2: Estimated Parameters

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Description</th>
<th>Unconstrained estimation</th>
<th>Constrained estimation</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\gamma$</td>
<td>Persistence of TFP growth</td>
<td>0.837</td>
<td>0.837</td>
</tr>
<tr>
<td>$\sigma_{\text{news}}$</td>
<td>Standard deviation of news shock</td>
<td>0.061 (0.018)</td>
<td>0.061 (0.018)</td>
</tr>
<tr>
<td>$\kappa_p$</td>
<td>Marginal cost slope of NKPC</td>
<td>1.000 (n.a.)</td>
<td>0.05 (n.a.)</td>
</tr>
<tr>
<td>$\omega_p$</td>
<td>Degree of price indexation</td>
<td>0 (n.a.)</td>
<td>0 (n.a.)</td>
</tr>
<tr>
<td>$\xi_w$</td>
<td>Frequency of wage adjustment</td>
<td>0.992 (0.006)</td>
<td>0.85 (n.a.)</td>
</tr>
<tr>
<td>$\omega_w$</td>
<td>Degree of wage reoptimization</td>
<td>0.695 (0.025)</td>
<td>1 (n.a.)</td>
</tr>
<tr>
<td>$b$</td>
<td>Habit persistence</td>
<td>0.880 (0.079)</td>
<td>0.461 (n.a.)</td>
</tr>
<tr>
<td>$\sigma_u$</td>
<td>Capital utilization parameter</td>
<td>0.006 (0.005)</td>
<td>0.0165 (0.030)</td>
</tr>
<tr>
<td>$S''$</td>
<td>Investment adjustment cost</td>
<td>0 (n.a.)</td>
<td>0 (n.a.)</td>
</tr>
<tr>
<td>$\rho$</td>
<td>Persistence of monetary policy</td>
<td>0 (n.a.)</td>
<td>0.710 (0.151)</td>
</tr>
<tr>
<td>$\theta_\pi$</td>
<td>Inflation response</td>
<td>2.920 (0.594)</td>
<td>3.000 (n.a.)</td>
</tr>
<tr>
<td>$\theta_{ygap}$</td>
<td>Output gap response</td>
<td>$-0.015$ (0.002)</td>
<td>$0.165$ (0.226)</td>
</tr>
<tr>
<td>$\theta_{\Delta y}$</td>
<td>Output growth response</td>
<td>2.130 (0.012)</td>
<td>0.553 (0.265)</td>
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<tr>
<td>$\lambda_{1,\pi}^{\text{news}}$</td>
<td>Risk loading on expected inflation</td>
<td>$-6.999$ (1.005)</td>
<td>$-11.165$ (3.390)</td>
</tr>
<tr>
<td>$\lambda_{1,\Delta u_c}^{\text{news}}$</td>
<td>Risk loading on expected change in marginal utility from consumption</td>
<td>$-15.032$ (1.771)</td>
<td>$-15.594$ (0.322)</td>
</tr>
</tbody>
</table>
Figure 1 Variance Decomposition of First Slope Shock
Figure 2 Impulse Response to First Slope Shock
Figure 3 Impulse Response to Contemporaneous TFP Shock

Figure 7: Impulse responses to contemporaneous TFP shock

Figure 8: Impulse responses to TFP news shock
Figure 4 Variance Decomposition to TFP News Shock

TFP

GDPinf

ffr

spread
Figure 5 Impulse Response to TFP News Shock

spread

TFP

c

GDPinf

ffr

long

news shock
slope shock

news shock
slope shock

39
Figure 6 Comparison of First Slope Shock and News Shock

Figure 7 Historical Simulations with TFP News Shock
Figure 8 Impulse Response to TFP News Shock
Figure 9 Impulse Response of Term Premium and Expectations Hypothesis to TFP News Shock
Figure 10 Impulse responses to TFP news shock for DSGE model (unconstrained estimates)
Figure 11 Impulse responses to TFP news shock for DSGE model (constrained estimates)