Quantitative Effects of Fiscal Foresight*

Eric M. Leeper† Alexander W. Richter‡ Todd B. Walker§

October 11, 2011

Abstract

Changes in fiscal policy typically entail two kinds of lags: the legislative lag—between when legislation is proposed and when it is signed into law—and the implementation lag—from when a new fiscal law is enacted to when it takes effect. These lags imply that substantial time evolves between when news arrives about fiscal changes and when the changes actually take place—time when households and firms can adjust their behavior. We identify two types of fiscal news—government spending and changes in tax policy. We identify news concerning taxes through the municipal bond market, and news concerning government spending through the Survey of Professional Forecasters. The main contribution of the paper is a mapping from reduced-form estimates of news into a DSGE framework. We find that news about fiscal policy is a time-varying process and show that ignoring the time variation can have important consequences in a conventional macroeconomic model.

*We would like to thank the editor, Roberto Perotti, and two anonymous referees for helpful comments on a previous draft. We would also like to thank participants at the 2010 NBER TAPES Conference, especially our discussants Stephanie Schmitt-Grohé and Ethan Ilzetzki.

†Indiana University, Monash University, and NBER, eleeper@indiana.edu

‡Indiana University, richtera@indiana.edu

§Indiana University, walkertb@indiana.edu

1 Introduction

Through a variety of not easily quantified sources—news reports, television, the internet, word-of-mouth—economic agents acquire foresight about future variables that are important to their decisions. Forward-looking decision-makers react to this news even before the variables are realized.

Much of the recent work on foresight focuses on news about future changes in technology,[1] but fiscal policy provides a more tangible example. Changes in fiscal policy typically entail two kinds of lags: the legislative lag—between when legislation is proposed and when it is signed into law—and the implementation lag—from when a new fiscal law is enacted to when it takes effect. These lags imply that substantial time evolves between when news arrives about fiscal changes and when the changes actually take place—time when households and firms can adjust their behavior. Although researchers have recognized that economic agents might change their behavior in anticipation of not-yet-realized tax changes [Hall (1971), Judd (1985), Branson et al. (1986), Poterba (1988), Sims (1988), Leeper (1989)], the theoretical and empirical implications of such foresight are only beginning to be studied [Yang (2005), Kriwoluzky (2009), Leeper et al. (2008, 2011), Mertens and Ravn (2008, 2011), Fisher and Peters (2010), Ramey (2011), Schmitt-Grohé and Uribe (2008)].

Leeper et al. (2011) and Leeper and Walker (2011) emphasize that the quantitative effects of foresight depend critically on the information processes governing the news. In principle, when the information flows are modeled "correctly" and then embedded into a dynamic stochastic general equilibrium (DSGE) model, it is possible to obtain accurate qualitative predictions of the effects of fiscal news (conditional on the DSGE model). Fiscal foresight and "news shocks," however, are generally difficult to pin down. The news processes embedded into a DSGE model must be imposed by the modeler and are therefore prone to misspecification. Leeper and Walker (2011) and Barsky and Sims (2011) emphasize that slight modifications to information processes governing foresight can lead to substantial changes in equilibrium outcomes.

Fiscal foresight creates special problems for structural VARs because it can produce equilibrium time series with a non-fundamental moving average component that misaligns the agents' and the econometrician's information sets [Ramey (2011), Leeper et al. (2008, 2011)]. Difficulties associated with non-fundamental moving average representations in macro models were first described by Hansen and Sargent (1980, 1991) and recently reiterated by Fernández-Villaverde et al. (2007). Economically meaningful shocks typically cannot be extracted from statistical innovations in conventional ways without making strong and unverifiable assumptions about information flows. Conventional econometric tools can yield false inferences by confounding shocks and incorrectly estimating dynamics. These difficulties suggest that one must be especially careful when examining foresight.

The primary contribution of this paper is to methodically construct news processes for fiscal policy—both taxes and spending—from data and map the news processes into a standard DSGE model. Following Fortune (1996), we identify news about tax policy changes through the use of municipal bonds (section 2.1). If asset markets are efficient, the yield spread between tax-exempt municipal bonds and treasury bonds should reflect the anticipated change

[1] For example, Beaudry and Portier (2006), Christiano et al. (2008) Jaimovich and Rebelo (2009), Comin et al. (2008), Fujiwara et al. (2008), and Leeper and Walker (2011).

in tax rates. We also identify news about changes in government spending following the approach described in Ramey (2009, 2011) (section 2.2). Ramey argues forcefully that at times significant changes in government spending are well anticipated. We use the Survey of Professional Forecasters to back out the amount of fiscal foresight contained in government spending. Section 3 lays out a DSGE model for policy evaluation and section 4 derives a unique mapping from the section 2 estimates of foresight to the DSGE model. The mapping equates the reduction in the variance of forecast errors attributable to foresight taken from empirical estimates to the DSGE model. We feed the two sources of fiscal policy news into a conventional new Keynesian model.

News about fiscal policy is a time-varying process. In some periods, like wars or significant tax reforms, agents have many quarters of foresight. Most of the time series consists of medium-to-low or no foresight. We quantify the impacts of fiscal foresight in Traum and Yang's (2010) new Keynesian model, which is a conventional model for policy analysis that has been fit to U.S. data. We augment this model with foresight and find that foresight can have both quantitative and qualitative effects on short- and medium-run dynamics—a result that is consistent with many papers in the literature. The typical assumption that news is a time-invariant process runs of the risk of *under-reporting* the impact of fiscal foresight because news periods get averaged over time. That is, the true effects of fiscal foresight may be masked by averaging high news periods with periods of no or low news. Using our calibration of news regimes, we address the extent to which a standard DSGE model estimated with time-invariant news may incorrectly imply that foresight is not relevant for explaining business cycle dynamics.

2 Identification of Fiscal Foresight

Recent papers have emphasized the difficulties that structural VARs may have in recovering the true impulse response function when agents have foresight [Leeper et al. (2011), Ramey (2011)]. Foresight generates equilibria in which the statistical shocks do not span the true information set of the agents. When estimating a DSGE model using likelihood or Bayesian methods, though, this problem no longer exists because the econometrician takes an explicit stand on the information set (conditional expectation) of the agent. For example, the modeler must specify the number of quarters agents have foresight and the extent to which agents have foresight.[2] But how does one go about calibrating the degree of foresight? Following Fortune (1996), we back out measures of foresight with respect to changes in tax policy using municipal bond market data. To identify foresight about government spending, we use data from the Survey of Professional Forecasters and Ramey (2011). We then show how to map these estimates into a DSGE model.

2.1 Identification of Tax Foresight

If markets are efficient, asset prices reflect all information currently available to market participants, especially news concerning the future paths of relevant variables. This hypothesis led Beaudry and Portier (2006) to include stock prices in a VAR in order to capture agents' expectations about future changes in productivity,

[2] One can even use this information to back out the true structural innovations from a VAR [Mertens and Ravn (2010)].

while Fisher and Peters (2010) use stock prices of government defense suppliers—which react to government defense purchases—to identify news about future government spending.

A more direct indicator is available for tax foresight: the preferential tax treatment of municipal bonds embeds the degree of tax foresight in certain financial variables.

In the United States, municipal bonds are exempt from federal taxes.[3] The differential treatment of municipal and treasury bonds has useful implications for identifying news about tax changes. If $\mathcal{Y}_t^M$ is the yield on a municipal bond at t and $\mathcal{Y}_t$ is the yield on a taxable bond, and assuming the bonds have the same term to maturity, callability, market risk, credit risk, and so forth, then an "implicit tax rate" is given by $\tau_t^I = 1 - \mathcal{Y}_t^M/\mathcal{Y}_t$. This is the tax rate at which the investor is indifferent between tax-exempt and taxable bonds. If participants in the municipal bond market are forward looking, the implicit tax rate should predict subsequent movements in individual tax rates. For example, if investors expect individual tax rates to rise (fall), they will demand higher (lower) yields on taxable bonds until they are indifferent between taxable and nontaxable bonds.

Several papers use event studies to document the ability of the municipal bond market to forecast changes in fiscal policy [Poterba (1986, 1989), Fortune (1996), Park (1997), Leeper et al. (2008)]. Using Yang's (2008) updated chronology of federal tax events, we estimate the response of implicit tax rates to major tax legislation that has taken place over the past half-century. The date of each tax event is set to the date of passage in the chamber of Congress that first passed the legislation, allowing us to evaluate how implicit tax rates are affected before the new policy is implemented.

Table 1 presents the results for bonds with maturity lengths of 1, 5, and 10 years, following the estimation strategy of Poterba (1986, 1989). Column 1 reports the predicted effect of each tax event on implicit tax rates. In general, tax events that reduce (increase) individual and corporate tax burdens were predicted to lower (raise) implicit tax rates, as the relative attractiveness of municipal bonds would fall (rise). The next three columns indicate whether the estimated effects of each tax event were statistically significant and/or matched their predicted sign. There are a total of 66 estimated coefficients based on 22 tax events over three maturity lengths. Of the 66 estimated coefficients, roughly three-quarters of the coefficients match their predicted sign, while two-fifths are statistically significant and match their predicted sign.

The table highlights an important feature of information from municipal bonds: the information content of implicit tax rates varies systematically across maturity lengths.[4] Over half of the estimated coefficients have the correct sign and are statistically significant for bonds with a 1-year maturity. At the 5-year horizon 11 events are significant and correctly signed, while only 4 are incorrectly signed and none are significant. Conversely, only five of the 22 tax events meet these criteria at 10-year horizons. Evidently, municipal bond yields are more informative about tax events in the near future (1 and 5 year horizons) than in the distant future. This is consistent with an inference that uncertainty about future tax policy and the impacts from contaminating factors (for example, call likelihood, credit risk) grow with the horizon being considered. Although event studies suffer from the drawback of requiring the

[3]Depending upon the type of bond, municipal bonds can also be exempt from the Alternative Minimum Tax, state, and local taxes. See Ang et al. (2010) for a thorough description of the municipal bond market.

[4]This characteristic, that the short-end of the municipal bond yield curve contains information about tax events while the long end does not, is well known, (e.g., Chalmers (1998)).

Table 1: Event Study Results*

Predicted Sign	Estimation Results† 1-yr	5-yr	10-yr	Date	Event‡
Positive	c,s	c	c	March 1954	Extended the 5 percentage point increase in corporate tax rates
Negative	c,s	c,s	c	Sept. 1963	Reduced individual and corporate income tax rates
Positive	c,s	c,s	c	Feb. 1968	Imposed individual and corporate income surtaxes
Positive	c,s	c	x	May 1969	Extended the 1968 surtaxes
Negative	c,s	c	c	Aug. 1969	Established the alternative minimum tax
Negative	c	c	x	Sept. 1971	Increased the personal exemption and standard tax deduction
Negative	c,s	x	x	Jan. 1975	Provided a 10% rebate on individual income taxes and reduced corporate income tax rates
Negative	c,s	c	c	Oct. 1975	Extended the 1975 tax reductions
Negative	x,s	x	x	Aug. 1976	Extended most of the 1975 tax law changes
Negative	c,s	c	c	March 1977	Increased standard deductions and further extended some of the 1975 tax law changes
Negative	x	c	x	Aug. 1978	Reduced individual and corporate income tax rates and raised the amount of capital gains that could be excluded from ordinary income taxes
Negative	x	x	c	July 1981	Reduced all individual income tax rates
Negative	c,s	c,s	c,s	July 1982	President Reagan vowed to not retreat from the 1981 tax cuts
Positive	c.s	c,s	c,s	March 1984	Effectively increased taxes by closing tax loopholes
Negative	c,s	c,s	c	Dec. 1985	House passage of Tax Reform Act of 1986: Dramatically reduced the number of tax brackets and repealed the capital gains exclusion and taxed them at the same rate as ordinary income
Negative	x	c,s	c,s	June 1986	Senate passage of Tax Reform Act of 1986
Negative	c	c,s	c,s	Sept. 1990	Increased individual income taxes, eliminated the "bubble" tax rate, and set a cap on the capital gains tax rate
Positive	c,s	c,s	c	May 1993	Created a new tax bracket for individual and corporate income taxes
Negative	c	c	c	June 1997	Reduced the top capital gains tax rate
Negative	c	x	x,s	May. 2001	Replaced the 5 existing individual tax brackets with 6 lower brackets
Negative	x	c,s	x	Apr. 2003	Accelerated the 2001 tax reductions
Negative	x	c,s	x	May 2004	Extended many of the provisions of the 2001 tax relief act
Negative	x,s	c,s	c,s	July 2004	Repealed the extraterritorial income exclusion

* Estimates are based on a feasible GLS procedure. Specifically, using homoscedastic OLS residuals, residual variances for each 24-month period were estimated and used to appropriately weight a second stage MA(1) regression with the change in the implicit tax rate as the dependent variable and each of the tax dummies as independent variables.
† A c denotes that the regression coefficient matches the predicted sign, an s denotes a regression coefficient that is statistically significant to a 95% confidence level, and an x denotes an incorrect regression coefficient sign.
‡ Unless otherwise noted, the date of each tax event was set to the date of passage in the chamber of Congress that first passed the legislation.

econometrician to specify the precise date(s) and relative importance of each event, at short maturities (under five years), they are suggestive of the information content of implicit tax rates.

With data on bond yields at various maturity lengths (see the data description in appendix A), it is possible to use the municipal bond yield curve as a measure of the expected *path* of tax rates. Implicit tax rates over two different maturity lengths yield a time series of implied forward tax rates [Poterba (1986) and Kochin and Parks (1988)]. Newly issued tax-exempt bonds with maturity T, a par value of \$1, and per-period coupon payments, C^M, will sell at par if

$$1 = \sum_{t=1}^{T} \frac{C^M}{(1+R_t^\tau)^t} + \frac{1}{(1+R_T^\tau)^T}, \tag{1}$$

where R_t^τ is the *after-tax* nominal interest rate for after-tax payments made in period t. No arbitrage conditions imply that a taxable bond with a similar maturity structure, paying

coupon, C, and selling at par will satisfy

$$1 = \sum_{t=1}^{T} \frac{C(1-\tau_t^e)}{(1+R_t^\tau)^t} + \frac{1}{(1+R_T^\tau)^T}, \tag{2}$$

where τ_t^e is the future tax rate expected to hold in period t.

If bonds sell at par, then the yield-to-maturity is equal to the coupon payments. Therefore, the implicit tax rate at time T is given by $\tau_T^I = 1 - C^M/C$. Subtracting (2) from (1) and solving for C^M/C gives

$$1 - \tau_T^I = \sum_{t=1}^{T} \omega_t (1-\tau_t^e), \tag{3}$$

where $\omega_t = \delta_t / \sum_{j=1}^{T} \delta_j$ and $\delta_t = (1+R_t^\tau)^{-t}$. Because the ω weights sum to unity, the implicit tax rate at T is the weighted average of discounted expected future tax rates over periods 1 to T. We can use this expression to back out the average expected future tax rate between periods s and t given by[5]

$$\tau_{s,t}^e \equiv \frac{\sum_{j=s+1}^{t} \delta_j \tau_j^e}{\sum_{j=s+1}^{t} \delta_j} = \frac{\tau_t^I \sum_{j=1}^{t} \delta_j - \tau_s^I \sum_{j=1}^{s} \delta_j}{\sum_{j=s+1}^{t} \delta_j}. \tag{4}$$

As described in Kochin and Parks (1988), the forward tax rate for the interval between periods s and t is a weighted average of the forward tax rates for that interval, with weights equal to the normalized discount factors for payments in that interval. In an environment with no change in tax policy and perfect information, we would expect these rates to be similar across maturity lengths.

Figures 1 and 2 plot the paths of one- and five-year forward tax rates from 1954 to 2005.[6] The shaded regions correspond to the total legislative lags, documented in Yang (2008). Over a short time horizon, where the likelihood of default and impact of callability is extremely low, substantial movements in the forward tax rates that occur within the shaded regions indicate that there is significant news about future tax policy that arrives before the legislation is passed. This news provides agents with some degree of tax foresight.

The Tax Reform Act of 1986 provides the clearest example of the information content of forward tax rates.[7] Over a one-year time horizon, the response is relatively small, since the policy was phased in over several years. However, five-year future tax rates correspond perfectly with the legislative lag, as the peak expectation coincides with the announcement of the policy and the trough expectation coincides with the implementation of the legislation. By the time the tax reform actually took effect, agents had factored the entire effect of the policy

[5] To derive this expression, rewrite (3) as $\tau_t^I = \sum_{j=1}^{T} = \omega_j \tau_j^e$. Then evaluate at time s and t and subtract to obtain $\sum_{j=s+1}^{t} \delta_j \tau_j^e = \tau_t^I \sum_{j=1}^{t} \delta_j - \theta_s^I \sum_{j=1}^{s} \delta_j$. Diving by the sum of the weights, $\sum_{j=s+1}^{t} \delta_j$, yields the expression for the average forward tax rate between periods s and t given in (4).

[6] Forward tax rates are computed using implicit tax rates for bonds with maturities lengths of 1 and 5 years. Note that 1 year forward tax rates are equivalent to implicit tax rates.

[7] This outcome is not surprising given Auerbach and Slemrod's (1997) evidence of how economic behavior adjusted during the long legislative and implementation processes associated with this act.

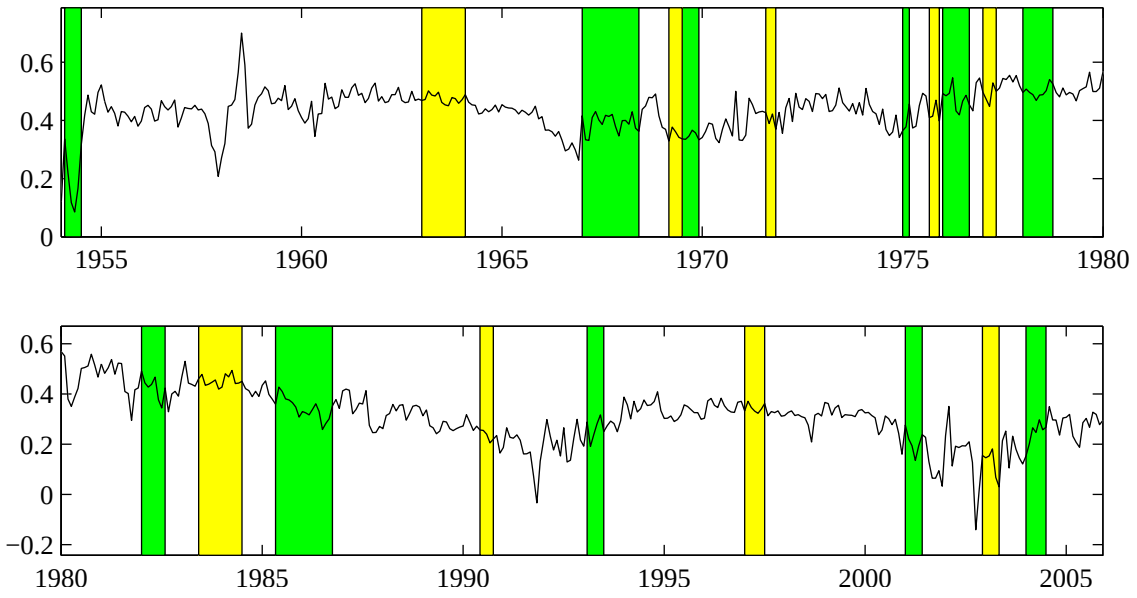


Figure 1: One-year forward tax rates ($t = 0$ to $t = 1$). Shaded regions correspond to tax events documented in Yang (2008). Shading differences only differentiate between events.

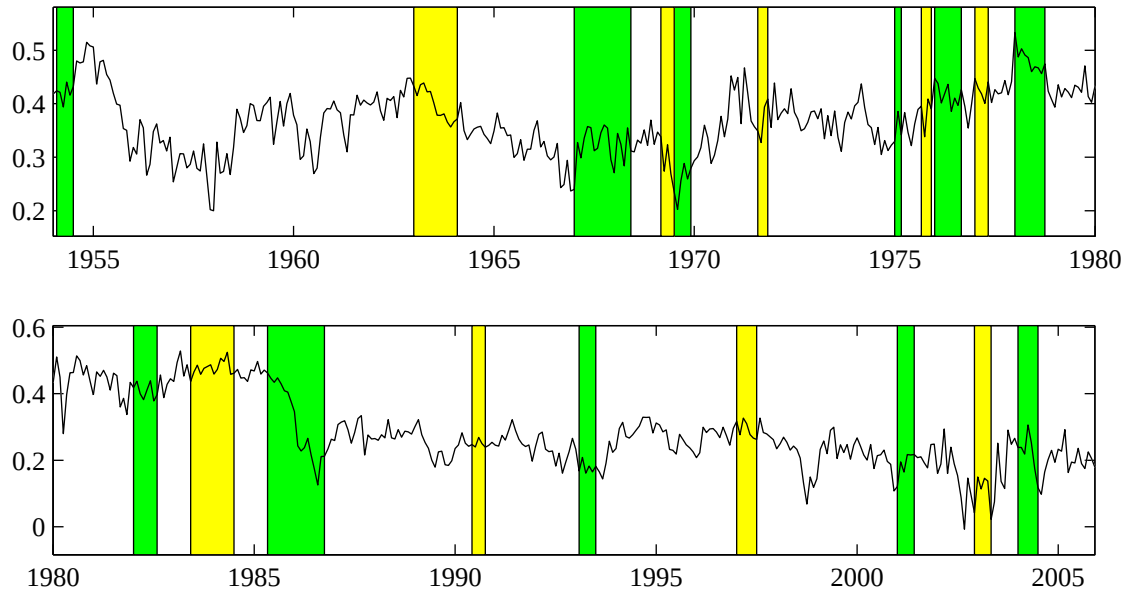


Figure 2: Five-year forward tax rates ($t = 1$ to $t = 5$). Shaded regions correspond to tax events documented in Yang (2008). Shading differences only differentiate between events.

into their expectations of taxes over the next five years. Although not all tax events are well aligned with agents' expectations, over shorter time horizons (under five years) forward tax rates are generally more responsive to proposed tax legislation than over longer time horizons.

One potential reason why forward tax rates do not correspond one-for-one with changes in tax policy is because risk must be taken into account when constructing the yield spreads between treasuries and municipal bonds. Differences in credit risk, call features, duration, underlying collateral, *etc.* all imply that investors would require a premium for holding municipal bonds.[8] To account for all unobservable risk that may arise, Fortune (1996) introduces a time-invariant "quality premium", ξ, in the relationship between yields on municipal bonds and treasuries. The risk-adjusted implicit tax rate is given by

$$\tau_t^{RI} = 1 - \frac{\mathcal{Y}_t^M + \xi}{\mathcal{Y}_t}. \tag{5}$$

When agents are compensated for holding the potentially risky municipal bonds, the yield spread between taxable treasury bonds and tax-exempt municipal bonds is reduced and the implicit tax rate falls.

To determine how well the risk-adjusted implicit tax rate forecasts changes in tax rates, we follow Fortune (1996) in constructing an ex-post tax rate. Let τ_{t+i} denote the representative agent's tax rate in period $t+i$. Given that coupons are typically paid semi-annually, we construct a series of future tax rates at a semi-annual frequency given by $\tau_{t+6}, \tau_{t+12}, \tau_{t+18}, ..., \tau_{t+6N}$, with t being the spot date and N the number of semiannual periods to maturity. The ex-post tax rates, given by

$$T_t = \sum_{i=1}^{N} \omega_i \tau_{t+6i},$$

are constructed from the known statutory tax rates over the period to maturity, where the weights are defined as above. Simply comparing the variances of implicit and ex-post tax rates would not reveal the information content of implicit tax rates because the risk premium may be correlated with implicit tax rates.

To more accurately determine how well municipal bonds forecast changes in tax rates, Fortune decomposes the ex-post tax rate into a convex combination of the risk-adjusted implicit tax rate, τ_t^{RI}, and the spot tax rate, τ_t, along with a forecast error to obtain

$$T_t = \alpha_1^\tau \tau_t^{RI} + (1 - \alpha_1^\tau)\tau_t + \varepsilon_t^\tau, \quad \varepsilon_t^\tau \sim N(0, \sigma_\tau^2). \tag{6}$$

The optimal weight given to each component depends on how much that component helps to predict changes in ex-post tax rates. Let $\zeta_{\tau^{RI}}$ denote the forecast error from predicting changes in the ex-post tax rate, conditional on the risk-adjusted implicit tax rate ($\zeta_{\tau^{RI}} = T - \tau^{RI}$). Let ζ_τ denote the forecast error from predicting changes in the ex-post tax rate, conditional on the spot tax rate alone ($\zeta_\tau = T - \tau$). The composite forecast error is given by the convex

[8] Using secured (defeased) municipal bonds, Chalmers (1998) finds that the most commonly mentioned contaminants of forward tax rates—default risk and callability—cannot explain why forward tax rates are poor measures of expected future tax rates, particularly at the long end of the yield curve. He concludes that fiscal uncertainty is the most likely explanation.

Table 2: Linear Regression Model with Fixed Coefficients*

$$T_t - \tau_t = \alpha_0^\tau + \alpha_1^\tau(\tau_t^I - \tau_t) + \alpha_2^\tau(1/\mathcal{Y}_t) + \alpha_3^\tau \text{TRA86} + \varepsilon_t^\tau$$

	1-Year 1965:1-1994:12			5-Year 1965:1-1990:12			10-Year 1965:1-1985:12		
	\$100K	\$75K	\$50K	\$100K	\$75K	\$50k	\$100K	\$75K	\$50K
α_0^τ	-0.000	-0.006	-0.007	-0.072	-0.070	-0.073	-0.046	-0.068	-0.077
$t_{\alpha_0^\tau}$	-0.036	-1.197	-1.153	-2.003	-2.748	-1.865	-1.072	-2.370	-1.677
α_1^τ	0.164	0.070	0.079	0.219	0.091	0.102	0.240	0.101	0.108
$t_{\alpha_1^\tau}$	3.792	3.110	3.010	4.899	3.461	3.049	4.622	3.137	2.621
α_2^τ	0.051	0.022	0.022	0.190	0.080	0.090	0.081	0.038	0.072
$t_{\alpha_2^\tau}$	1.549	1.180	1.060	1.341	1.030	0.906	0.354	0.293	0.427
α_3^τ	0.002	0.003	0.001	0.077	0.055	0.058	0.080	0.058	0.062
$t_{\alpha_3^\tau}$	0.169	0.446	0.071	3.894	4.930	4.075	3.753	4.706	3.864
DW	1.777	2.058	1.964	2.054	2.387	2.193	2.037	2.395	2.186
Q_{12}	15.164	9.781	3.715	11.750	21.285	9.867	10.065	25.688	9.873
	(0.233)	(0.635)	(0.988)	(0.466)	(0.046)	(0.628)	(0.610)	(0.012)	(0.627)

* Cochrane-Orcutt estimation was used to correct for serial correlation. The Box-Ljung statistic (Q_{12}) tests for serial correlation over a 12-quarter period. The corresponding p-value is in parentheses. The correction was successful in all but two cases.

combination of the two, $\varepsilon^\tau = \alpha_1^\tau \zeta_{\tau^{RI}} + (1 - \alpha_1^\tau)\zeta_\tau$. The optimal weight, α_1^τ, is chosen to minimize the variance of the forecast error. This weight is given by

$$\alpha_1^\tau = \frac{\sigma^2_{\zeta_\tau}}{\sigma^2_{\zeta_\tau} + \sigma^2_{\zeta_{\tau^{RI}}}}, \tag{7}$$

where $\sigma^2_{\zeta_\tau}$ and $\sigma^2_{\zeta_{\tau^{RI}}}$ are the variances of the forecast errors ζ_τ and $\zeta_{\tau^{RI}}$. More weight is given to the variable that has the smaller forecast error variance. For example, if agents have perfect foresight (that is, if agents know exactly what their tax rates are going to be for N semiannual periods) and markets are efficient, the variance of the forecast error conditional on the risk-adjusted implicit tax rate, $\sigma^2_{\zeta_{\tau^{RI}}}$, would be zero and α_1^τ would be set to unity.

Substituting (5) into (6) and re-arranging gives

$$T_t - \tau_t = \alpha_1^\tau(\tau_t^I - \tau_t) + \alpha_2^\tau(1/\mathcal{Y}_t) + \varepsilon_t^\tau, \tag{8}$$

where α_1^τ measures the information content of municipal bonds and $\alpha_2^\tau = -\alpha_1^\tau \xi$ measures the risk premium. We can now disentangle the effects of risk to back out the informational content of implicit tax rates.

Table 2 displays the results of the estimation of (8) using marginal income tax rates for married individuals filing joint returns collected from Internal Revenue Service publications and the Tax Policy Center. The series of actual and ex-post tax rates were constructed using marginal tax rates for investors earning \$100,000, \$75,000, and \$50,000 annually in constant 1980 dollars. The yields to maturity are taken from tax-exempt prime-grade general-obligation municipal bonds obtained from Salomon Brothers' Analytical Record of Yields and Yield Spreads for maturity lengths of 1, 5 and 10 years.[9] As the table reports, the information

[9] Following Fortune (1996), we also include a dummy variable for the 1986 Tax Reform Act (TRA). This dummy variable is included to account for the significant change in the market structure of the municipal bond market caused by the TRA.

Table 3: Linear Regression Model with Variable Coefficients*

$$T_t - \tau_t = \alpha^\tau_{0,t} + \alpha^\tau_{1,t}(\tau^I_t - \tau_t) + \alpha^\tau_{2,t}(1/\mathcal{Y}_t) + \alpha^\tau_{3,t}\text{TRA86} + \varepsilon^\tau_t$$
$$\alpha^\tau_{j,t} = \alpha^\tau_{j,t-1} + \eta^\tau_{j,t} \text{ (j=1,2,3)} \quad \eta^\tau_t \sim N(0, \Delta^2_\tau) \quad \varepsilon_t \sim N(0, \sigma^2_\tau)$$

	1-Year 1965:1-1994:12			5-Year 1965:1-1990:12			10-Year 1965:1-1985:12		
	\$100K	\$75K	\$50K	\$100K	\$75K	\$50k	\$100K	\$75K	\$50K
$\delta_{0,\tau}$	0.0001	0.1468	0.0019	0.0000	0.2199	0.0097	0.0001	0.2592	0.0014
$\delta_{1,\tau}$	0.6311	0.8500	0.3797	0.6058	1.0004	1.2521	0.2542	2.0762	2.6473
$\delta_{2,\tau}$	0.1232	0.0007	0.0001	0.0005	0.0028	0.0008	0.0006	0.0058	6.0488
$\delta_{3,\tau}$	0.0022	0.0020	0.0023	0.0123	0.0061	0.0068	0.0160	0.0099	0.0087
σ_τ	0.0180	0.0117	0.0084	0.0328	0.0416	0.0245	0.0246	0.0255	0.0258

* Maximum Likelihood estimation was used to obtain the standard errors of the transition equation steps (square roots of the diagonal elements in Δ^2_τ), $\delta_{0,\tau}$, $\delta_{1,\tau}$, $\delta_{2,\tau}$, $\delta_{3,\tau}$. All off diagonal entries are assumed to be zero. The parameter σ_τ is the standard deviation of the measurement equation.

parameter, α^τ_1, is of the correct sign and statistically significant for all maturity lengths and income groups, suggesting that the information parameter contains relevant news about future tax rates. Not surprisingly, the information content of implicit tax rates is greatest for agents who face higher marginal tax rates. The risk premium parameter, α^τ_2, is also positive across most maturity lengths, but not with statistical significance. This reflects that municipal bonds pose little risk to investors over a short horizon.

To capture the time varying nature of the information content contained in municipal bonds and allow for time-varying risk premia, Fortune (1996) estimates a version of (8) where the coefficients vary with time according to a random walk specification given by

$$\alpha^\tau_{j,t} = \alpha^\tau_{j,t-1} + \eta^\tau_{j,t}, \qquad j = 1, 2, \qquad \eta^\tau_{j,t} \sim N(0, \delta^2_{j,\tau}). \tag{9}$$

The standard deviations of the information parameter and risk premium give an indication of the amount of time variation in these parameters. Equations (8) and (9) form a state-space representation for which the Kalman filter can be used to estimate the model.[10]

Table 3 reports the estimation allowing for time-varying parameter values. Notice that the standard deviation is largest for the information parameter ($\delta_{1,\tau}$). This suggests that the information content of municipal bonds, and hence foresight with respect to tax policy, is very much a time-varying process. Figure 3, which plots the predicted path of the information parameter based on the marginal tax rate for an individual earning \$75,000 in constant 1980 dollars, also demonstrates this point. For the decades of the 1970s and 1990s, the information contained in municipal bonds is negligible relative to the 1980s. The spikes in the information parameter correspond to the Economic Recovery Tax Act of 1981 and the Tax Reform Act of 1986.

2.2 Identification of Government Spending Foresight

To identify foresight with respect to government spending, we follow Ramey (2011) in using the Survey of Professional Forecasters (SPF) conducted by the Federal Reserve Bank of Philadelphia. The data we examine are mean *forecasts* of real federal government consumption and gross investment from 1981Q3 to 2010Q1 over horizons ranging from one to four quarters.[11] Data on quarterly nomi-

[10] See Durbin et al. (2004) for more details on the estimation procedure.

[11] For in-depth analysis on the explanatory power of the SPF see Perotti (2011).

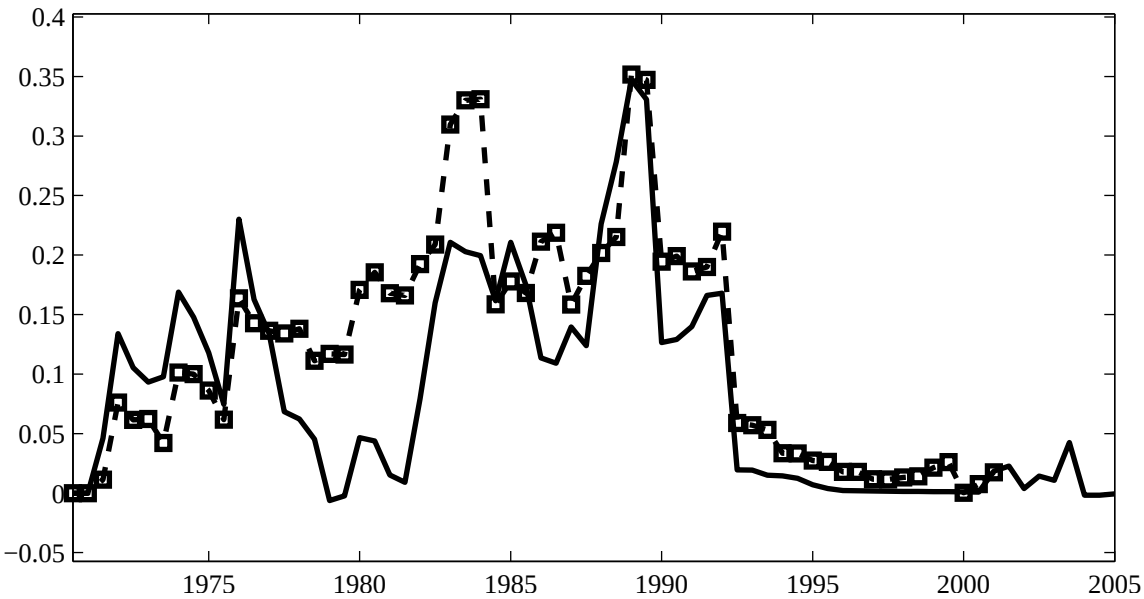


Figure 3: Estimated path of the time-varying information parameter α_1^τ. The solid black and dotted-dashed lines correspond to bonds with maturity lengths of 1 and 5 years. Estimation is based on the marginal tax rate for an individual earning $75,000 in constant 1980 dollars.

nal federal government consumption and gross investment spending over the same period come from the National Income and Product Accounts, published by the Bureau of Economic Analysis (BEA). Federal government consumption expenditures and gross investment in chained 2005 dollars was generated using the component-specific real GDP quantity index and annual component-specific nominal GDP. Appendix A contains a complete description of the data.

Ramey (2011) (and references therein) provides ample empirical evidence for foresight with respect to government spending. Among other tests, she finds that one- and four-quarter ahead professional forecasts Granger cause VAR shocks. Using data from 1939 to 2008, she also finds that a "defense news" variable corresponding to major war dates has significant explanatory power in forecasting changes in government and defense spending. Figure 4 plots real federal government spending along with Ramey's war dates. As is evident from this picture, defense news is often followed by stark changes in government consumption and investment expenditures.

Analogous to the analysis for tax foresight, we assume that forecasts of government spending can be decomposed into two components,

$$G_{t+j} = \alpha_{j,t}^G G_{t+j|t} + (1-\alpha_{j,t}^G)\rho_G^j G_t + \varepsilon_t^G, \quad j = 1,\ldots,4, \quad \varepsilon_t^G \sim N(0,\sigma_G^2) \tag{10}$$

$$\alpha_{j,t}^G = \alpha_{j,t-1}^G + \eta_{j,t}^G, \qquad \eta_{j,t}^G \sim N(0,\delta_{j,G}^2). \tag{11}$$

The first component, $G_{t+j|t}$, is the SPF forecast of government spending at time $t+j$ conditional on time t information. We utilize SPF forecasts of real government spending, which range from one to five quarters. The second component assumes an AR(1) process for government spending.[12] We fit the AR(1) model using OLS on quarterly first-differenced real federal

[12]Given that most DSGE models specify an AR(1) structure for government, we thought that the most

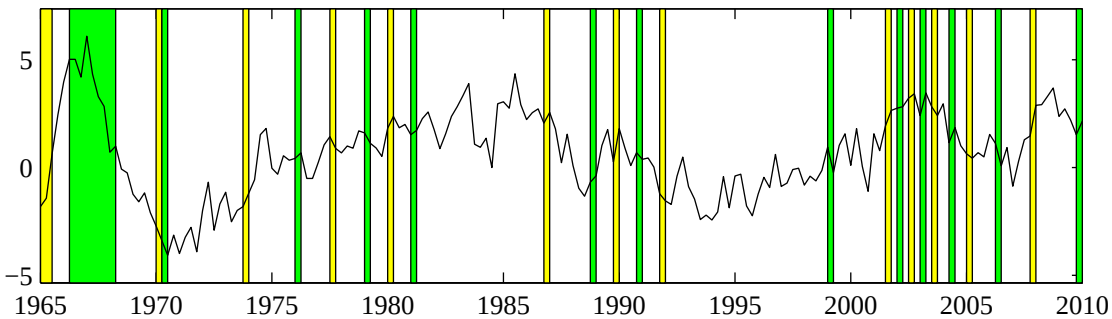


Figure 4: Annual log deviations in real federal government expenditures. Shaded regions correspond to defense spending events documented in Ramey (2009). Note that shading differences are only intended to help differentiate between events.

government expenditures from 1981Q3 to 2010Q1. Analogous to the tax foresight case and (7), $\alpha^G_{j,t}$ is determined by whichever forecast has the smaller forecast error variance.

To foreshadow results somewhat, section 4 will interpret "news" about government spending as coming from the SPF forecasts only, and derive a mapping between $\alpha^G_{j,t}$ and the agents' information set. This implies that the innovation, ε^G_t, in (10) is also an innovation with respect to the agents' information sets. But, as (10) makes clear, what is relevant is how much of the agents' innovation is due to news ($\alpha^G_{j,t}$) and how much can be attributed as coming from conditioning on past observations of government spending ($1-\alpha^G_{j,t}$). In other words, the error term from the reduced-form regression can be interpreted as "pure news" if $\alpha^G_{j,t}=1$ and "no news" if $\alpha^G_{j,t}=0$. In this sense we allow the data to determine the information flows available to agents.

As with tax foresight, the information parameter for government spending may be a time-varying process and we use a Kalman filter to back out the path of the information parameter. Figure 5 plots the $\alpha^G_{j,t}$ parameter for $j=1,2,3,4$ from 1982 through 2009.[13] The estimation reveals that news about government spending is also a time-varying process. The increase in the information parameter throughout the decade of 2000 is consistent with the increase in the frequency of defense spending events documented by Ramey, figure 4.[14]

3 Model for Policy Evaluation

We adopt a conventional new Keynesian (NK) model based on Traum and Yang (2010), which incorporates several frictions that have become standard in the literature. We provide a quick overview of the model and direct interested readers to Traum and Yang (2010) for a more thorough description.

The model includes two types of households: savers, denoted by S and in proportion μ,

appropriate horse race was to compare the SPF to the forecasts from an AR(1). More elaborate time series specifications for the government spending process were estimated but model selection criteria (e.g., AIC) preferred the AR(1) specification.

[13]For the sake of brevity, we do not report the variance in the time variation but it is available upon request from the authors.

[14]Our estimates of α^G_t included up to a 5-quarter horizon but showed little informational content and are, therefore, omitted from figure 5.

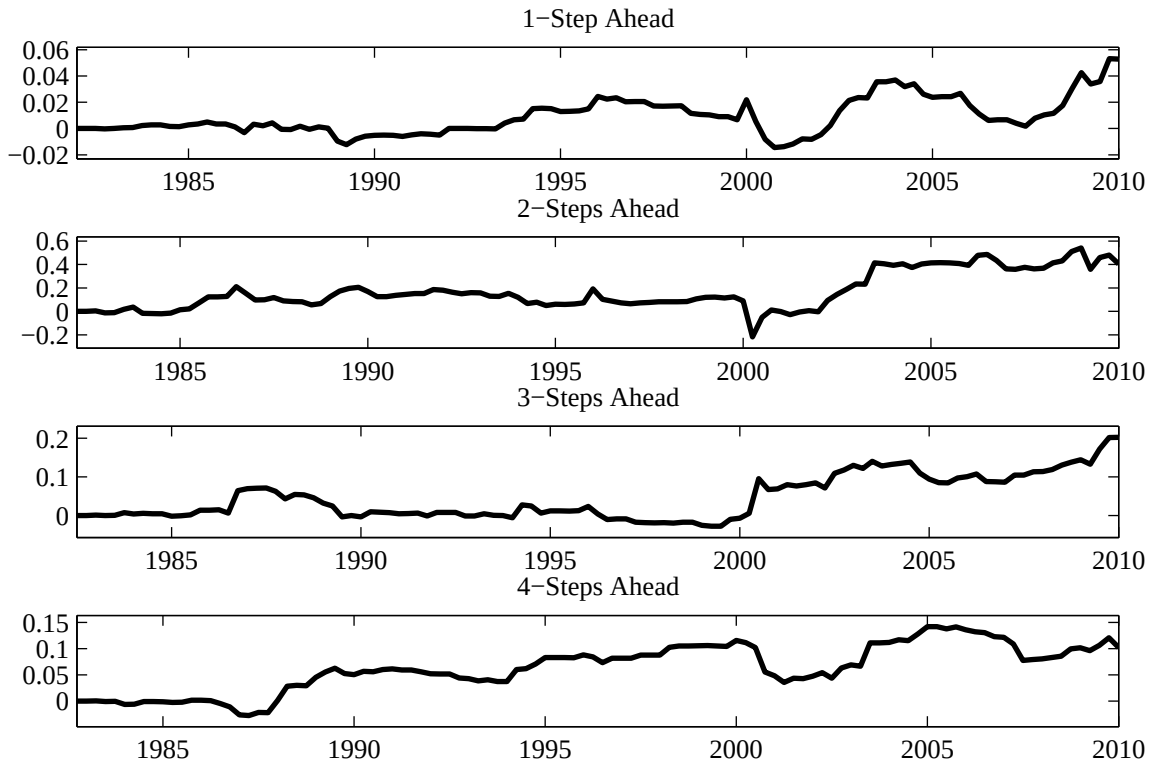


Figure 5: Time-varying information parameter, α_1^G, for 1, 2, 3, and 4-step-ahead forecasts of real government spending from 1982-2009.

who have access to a complete set of contingent claims, and non-savers, denoted by N and in proportion $(1-\mu)$, who each period consume their entire disposable income. The continuum of agents have common preferences, as represented by those of agent $j \in [0,1]$

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{c_t^A(j)^{1-\gamma} - 1}{1-\gamma} - \frac{L_t^A(j)^{1+\kappa}}{1+\kappa} \right] \tag{12}$$

for $A \in \{S, N\}$, where $0 < \beta < 1$ is the household's discount rate, $\gamma \geq 0$ is the constant of relative risk-aversion, and $\kappa \geq 0$ is the inverse of the Frisch labor supply elasticity. $c_t^A(j)$ and $L_t^A(j)$ are, respectively, consumption of the final good and the quantity of labor supplied at time t by agent j. Each individual agent's labor input, $\ell \in [0,1]$, is supplied in a monopolistically competitive setting.

The budget constraint for saver $j \in (0, 1-\mu)$ is

$$H_t(j) + (1-\tau_t^K)\frac{R_t^K v_t(j) k_{t-1}(j)}{P_t} + \frac{R_{t-1} b_{t-1}(j)}{\pi_t} = c_t^S(j) + \frac{i_t(j)}{1+\tau_t^C} + b_t(j), \tag{13}$$

where $b_t(j)$ and $k_t(j)$ denote the level of nominal riskless government bonds and the stock of capital carried into period $t+1$, P_t is the after-tax consumer price level, R_t and $\pi_t = P_t/P_{t-1}$ are the gross nominal interest rate on bonds purchased at time t and the gross inflation rate, and τ_t^L, τ_t^K, and τ_t^C are taxes levied against labor income, the return on capital, and

consumption. The presence of consumption taxes distinguishes the producer price index, $\bar{P}$, from the consumer price index, $P_t = (1+\tau_t^C)\bar{P}_t$. The term $H_t(j)$ represents individual j's human wealth (net labor income) and is given by $H_t(j) \equiv (1-\tau_t^L)\int_0^l \frac{W_t(\ell)}{P_t}\ell_t^S(j,\ell)d\ell + z_t(j) + d_t(j)$, where $W_t(\ell)$ is the nominal wage for labor type ℓ, $z_t(j)$ are government transfers, and $d_t(j)$ denotes the share of nominal firm profits received in the form of dividends by agent j. The law of motion for capital is given by

$$k_t(j) = (1-\delta[v_t(j)])k_{t-1}(j) + \left[1 - s\left(\frac{i_t(j)}{i_{t-1}(j)}\right)\right] i_t(j), \tag{14}$$

where $s(\cdot)$ is the investment adjustment cost function that satisfies the properties $s(1) = s'(1) = 0$ and $s''(1) \equiv s > 0$. The depreciation rate, δ, is positively related to the utilization rate, v_t, and is given by $\delta[v_t(j)] = \delta_0 + \delta_1(v_t(j)-1) + \frac{\delta_2}{2}(v_t(j)-1)^2$, where δ_0, δ_1, and δ_2 are calibrated parameters.[15]

The budget constraint for non-saver $j \in (1-\mu, 1]$, who does not have access to asset markets, is $c_t^N(j) = (1-\tau_t^L)\int_0^1 \frac{W_t(\ell)}{P_t}\ell_t^N(j,\ell)d\ell + z_t(j)$.

The total supply of labor services by savers and non-savers is identical. Specifically, $L_t^S(j) = L_t^N(j) = \int_0^1 \ell_t(\ell)d\ell \equiv L_t$. A labor clearinghouse purchases the differentiated labor inputs and groups them to generate a composite labor service, L_t, according to CES technology, $L_t = [\int_0^1 l_t(l)^{1/(1+\eta^w)}dl]^{1+\eta^w}$, where η^w denotes an exogenous markup to wages. Maximizing profits for a given level of labor yields firm i's demand function for a particular labor input given by $l_t(l) = L_t^d(W_t(l)/W_t)^{-(1+\eta^w)/\eta^w}$, where L_t^d represents the demand for composite labor services and $\psi^w \equiv (1+\eta^w)/\eta^w$ is the elasticity of substitution between inputs.

The production sector consists of monopolistically competitive intermediate goods producing firms who produce a continuum of differentiated inputs and a representative final goods producing firm. Each firm $i \in [0,1]$ in the intermediate goods sector produces a differentiated good, $y_t(i)$, with identical technologies given by $y_t(i) = (v_t k_{t-1}(i))^\alpha(\ell_t(i))^{1-\alpha}(K_{t-1}^G)^{\alpha^G}$, where $k_t(i)$ and $\ell_t(i)$ denote the capital stock and level of employment used by firm i.

The representative final goods producing firm purchases inputs from the intermediate goods producing firms to produce a composite good, Y_t, according CES technology, $Y_t = [\int_0^1 y_t(i)^{1/(1+\eta^p)}di]^{1+\eta^p}$, where η^p denotes an exogenous markup to the intermediate goods' prices. Analogous to the labor market, firm i's demand function for intermediate inputs is given by $y_t(i) = Y_t(\bar{p}_t(i)/\bar{P})^{-(1+\eta^p)/\eta^p}$, where $\bar{p}_t$ is the price of intermediate good i and $\bar{P}_t$ is the price of the final good.

Both wages and prices adjust according to a Calvo pricing mechanism. Each intermediate goods producing firm may reset its price only with probability $(1-\omega_p)$. Firms that are able to make optimal adjustments to their price level choose their price level, $\bar{p}_t(i)$, to maximize the sum of discounted future profits. Similarly, a union has the opportunity to adjust the nominal wage rate with probability $(1-\omega_w)$. It freely adjusts the nominal wage rate by choosing the optimal wage rate, $\widetilde{W}_t(\ell)$, to maximize the lifetime utility of households.

The fiscal authority finances government consumption, G_t, government investment, G_t^I, and government transfers, Z_t, through proportional taxes levied against consumption, labor

[15] δ_1 is calibrated so that $v = 1$ in steady-state. The parameter $\psi \in [0,1)$ is defined so that $\delta''(1)/\delta'(1) = \delta_2/\delta_1 \equiv \psi/(1-\psi)$.

income, and capital returns and by issuing one-period nominal debt. The government's flow budget constraint is

$$B_t + \tau_t^K \frac{R_t^K}{P_t} v_t K_{t-1} + \tau_t^L \frac{W_t}{P_t} L_t + \frac{\tau_t^C}{1+\tau_t^C} C_t = \frac{R_{t-1} B_{t-1}}{\pi_t} + G_t + G_t^I + Z_t. \tag{15}$$

Productive government capital evolves according to $K_t^G = (1-\delta^G)K_{t-1}^G + G_t^I$. Fiscal variables are determined by the following rules

$$\hat{\tau}_t^K = \rho_K \hat{\tau}_{t-1}^K + (1-\rho_K)\left(\varphi_K \hat{Y}_t + \gamma_K \hat{s}_{t-1}^b\right) + \phi_{KL}\sigma_L \varepsilon_t^L + \sum_{i=0}^{q} \theta_i^K \varepsilon_{t-i}^K \tag{16}$$

$$\hat{\tau}_t^L = \rho_L \hat{\tau}_{t-1}^L + (1-\rho_L)\left(\varphi_L \hat{Y}_t + \gamma_L \hat{s}_{t-1}^b\right) + \phi_{KL}\sigma_K \varepsilon_t^K + \sum_{i=0}^{q} \theta_i^L \varepsilon_{t-i}^L \tag{17}$$

$$\hat{G}_t = \rho_G \hat{G}_{t-1} - (1-\rho_G)\gamma_G \hat{s}_{t-1}^b + \sum_{i=0}^{q} \theta_i^G \varepsilon_{t-i}^G \tag{18}$$

$$\hat{G}_t^I = \rho_{GI} \hat{G}_{t-1}^I - (1-\rho_{GI})\gamma_{GI} \hat{s}_{t-1}^b + \sigma_{GI}\varepsilon_t^{GI} \tag{19}$$

$$\hat{Z}_t = \rho_Z \hat{Z}_{t-1} - (1-\rho_Z)\gamma_Z \hat{s}_{t-1}^b + \sigma_Z \varepsilon_t^Z \tag{20}$$

$$\hat{\tau}_t^C = \rho_C \hat{\tau}_{t-1}^C + \sigma_C \varepsilon_t^C, \tag{21}$$

where $s_{t-1}^b \equiv B_{t-1}/Y_{t-1}$ and $\varepsilon_t^s \sim i.i.d.\ N(0,1)$ for $s \in \{K, L, GC, GI, C, Z\}$.

Only consumption taxes are exogenous. In the United States, such taxes are relatively unimportant and do not seem to co-move with other variables [Leeper et al. (2010a)]. The remaining five fiscal instruments respond systematically to the debt-output ratio, $\hat{s}_{t-1}^b$, to stabilize debt. Capital and labor taxes also include automatic stabilizers that react to output fluctuations. Those tax rates and government consumption include both unanticipated and anticipated "shocks." Anticipated shocks, denoted by $\sum_{i=0}^{q} \theta_i^j \varepsilon_{t-i}^j$ for $j \in \{K, L, G\}$, reflect fiscal foresight. Finally, consistent with actual U.S. tax behavior, capital and labor taxes permit correlation through the parameter ϕ_{KL}.

The monetary authority sets interest rate policy according to a Taylor-type rule given by

$$\hat{R}_t = \rho_r \hat{R}_{t-1} + (1-\rho_r)\left[\phi_\pi \hat{\pi}_t + \phi_y \hat{Y}_t\right], \tag{22}$$

so that the nominal interest rate responds to fluctuations in both output and inflation.

4 Mapping of News into DSGE Models

Recent work that introduces news shocks into DSGE models must take a specific stance on the news process [for example, Christiano et al. (2008), Schmitt-Grohé and Uribe (2008), Fujiwara et al. (2008), Mertens and Ravn (2011)]. As emphasized in Leeper and Walker (2011) and Barsky and Sims (2011), the assumed information process can have profound effects on equilibrium dynamics; assuming agents have one quarter of foresight will lead to vastly different conclusions than an estimated process that allows eight quarters of foresight. We now turn to mapping estimates of news from section 2 into the model in section 3.

There are two dimensions to fiscal foresight—horizon and intensity. The foresight horizon measures how far in advance agents are aware of potential changes to fiscal policy. Foresight intensity measures how confident agents are about pending changes to fiscal variables. We now describe how these dimensions are calibrated.

4.1 Foresight Horizon The information parameters, α_1^τ and α_1^G, determine the extent to which news about taxes and government spending varies with time. While we are not conducting formal econometric tests of structural breaks in the time series of the information parameters, it is evident from figures 3 and 5 that there is substantial time variation in the news content of municipal bonds and the SPF, and the high news (high α_1) periods correspond nicely with historical episodes of fiscal changes.

We calibrate our foresight *horizon* to match specific periods in U.S. data allowing for "high", "medium" and "low" news periods for taxes and "high" and "medium" news periods for government spending. We calibrate the high degree of tax foresight specification using data from the 1980s, a high-news decade because of two major changes to the tax code—the Economic Recovery Tax Act of 1981 (HR 4242) and the Tax Reform Act of 1986 (HR 4170). Both bills implemented major changes to the tax code and had an average lag (from first announcement by the president to when the tax change took effect, including phased-in tax changes) of well over two years [Yang (2008)]. The lags associated with tax changes in the 1980s suggest a forecast horizon of at least eight quarters. We therefore assume the "high-news" tax regime has a foresight horizon of eight quarters ($q = 8$).

The medium and low degrees of tax foresight are calibrated to match the data from the 1970s and 1990s, respectively. There were several changes to the tax code in the 1970s—Revenue Act of 1971, Tax Reduction Act of 1975, Revenue Adjustment Act of 1975, Tax Reform Act of 1976, Tax Reduction and Simplification Act of 1977 and the Revenue Act of 1978. Most of these were relatively minor compared to the Tax Reform Act of 1986 and had significantly smaller lag times (under one year on average). This is confirmed by figure 3, which shows the information content of municipal bonds was, on average, smaller than for the 1980s. Conversely, the information contained in municipal bonds from 1990 through 2001 is nearly zero. For the medium-news regime, we assume agents have four quarters of foresight ($q = 4$) and for the low-news regime, we assume agents have only two quarters of foresight ($q = 2$). Again, these specifications match the lags for the major tax changes over these two decades as recorded by Yang (2008).

For government spending foresight, we use two specifications of news—high and medium. The high-news period is calibrated to match the data from 2000 through 2009. As shown in figure 5, the information content of the SPF's forecasts of changes in government spending over 1-, 2-, 3-, and 4-quarter horizons is highest during this decade. This corresponds nicely with Ramey's (2009) narrative in figure 4. The 2000s contained many defense spending increases: [i] In 2002Q1, the Bush administration calls for an increase in the Pentagon budget over the next 5 years; [ii] In 2002Q3, there were announced increases in the Department of Defense budget over the next 10 years to deal with counter-terrorism efforts and the response to 9/11; [iii] Several increases in spending to finance the wars in Afghanistan and Iraq. For the "high-news" regime, we allow for a forecast horizon of four quarters ($q = 4$). This is justified because estimates of the the information parameter, α_1^G, for the one- through four-step-ahead forecasts of real government spending from 2000 through 2009 are all significantly different

from zero. The medium degree of foresight is calibrated to match data from 1980 through 2000. The functional form of the government spending process assumes three quarters of foresight ($q = 3$), which is less than the maximum provided by the SPF of five. We specify only three quarters of foresight because the four- and five-step-ahead forecasts where given nearly zero weight in the estimation of (10) during these decades.

4.2 Foresight Intensity

Specifying the foresight horizon delivers the functional form of the tax and government spending processes (MA(q)), but we also need to calibrate the values of the moving-average coefficients in these processes. We now describe the mapping from the reduced-form estimates in section 2 to the moving-average coefficients in the tax (capital and labor) and government spending rules [see (16), (17), and (18)].

To illustrate of the mapping, consider the following moving-average representation for tax rates

$$\tau_t = \varepsilon_{t-1} - \theta\varepsilon_t. \tag{23}$$

If $|\theta| < 1$, then (23) is a non-fundamental moving-average representation, and the space spanned by current and past tax rates, $\{\tau_{t-j}\}_{j=0}^{\infty}$, is smaller than the space spanned by the structural innovations, $\{\varepsilon_{t-j}\}_{j=0}^{\infty}$.[16]

One consequence of this result is that the variance of the one-step-ahead forecast error for agents conditioning on structural innovations is smaller than the forecast error variance for agents conditioning *only* on current and past tax rates. To show this analytically, we must derive the Wold representation of (23), which comes from flipping the root, θ, outside of the unit circle

$$\tau_t = \tilde{\varepsilon}_t - \theta\tilde{\varepsilon}_{t-1} \tag{24}$$

$$\tilde{\varepsilon}_t = \left[\frac{L-\theta}{1-\theta L}\right]\varepsilon_t. \tag{25}$$

Representation (24) is the Wold representation where $\tilde{\varepsilon}_t$ is the one-step-ahead forecast error associated with forecasting τ_t conditional on $\{\tau_{t-j}\}_{j=1}^{\infty}$. This representation shows that current and past τ_t span an equivalent space to current and past $\tilde{\varepsilon}_t$, which is a strictly smaller space than ε_t. To prove this, note that using the Wiener-Kolmogorov optimal prediction formula yields the variance of the one-step-ahead forecast error using representation (24)

$$\begin{aligned} E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2 &= E\{(L-\theta)\varepsilon_{t+1} - L^{-1}[(1-\theta L)-1]\tilde{\varepsilon}_t\}^2 \\ &= E\{(L-\theta)\varepsilon_{t+1} - (L-\theta)\varepsilon_{t+1} + \tilde{\varepsilon}_{t+1}\}^2 \\ &= \text{var}(\tilde{\varepsilon}_{t+1}) = \sigma_\varepsilon^2, \end{aligned} \tag{26}$$

where the last equality follows because the term, $\frac{L-\theta}{1-\theta L}$, known as a Blaschke factor, has a covariance generating function of one (see Lippi and Reichlin (1994)) and hence var(ε_t)=var($\tilde{\varepsilon}$) $= \sigma_\varepsilon^2$.

[16] Other papers have assumed an i.i.d. process for news (e.g., $\tau_t = \varepsilon_{1,t-1} - \theta\varepsilon_{2,t}$, where $\varepsilon_{1,t-1}$ and $\varepsilon_{2,t}$ are orthogonal at all leads and lags). Given that the mapping requires a matching of forecast error variances at different horizons, the mapping derived in this section can easily extend to models with i.i.d. news processes.

Suppose now that agents are able to condition on current and past structural innovations directly. These agents are able to use (23) to forecast next period's tax rate. The variance of the forecast error for this process is given by

$$\begin{aligned} E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^{\infty}]\}^2 &= E\{(L-\theta)\varepsilon_{t+1} - L^{-1}[(L-\theta)+\theta]\varepsilon_t\}^2 \\ &= \theta^2 E\{\varepsilon_{t+1}\}^2 = \theta^2\sigma_\varepsilon^2. \end{aligned} \tag{27}$$

Comparing this forecast error variance with (26) shows that the moving-average coefficient, θ, determines the degree to which agents conditioning on the structural shocks are better informed. As $\theta \to 0$, agents who observe the structural innovations have perfect one-step-ahead foresight in the sense that they observe $\varepsilon_t = \tau_{t+1}$ and the corresponding forecast error is zero. As $\theta \to 1$, the information sets and the variance of forecast errors converge. Calibrating θ also calibrates agents' foresight intensity.

Recall from section 2 that the contemporaneous risk-adjusted implicit tax rate, τ_t^{RI}, is the weighted sum of *future* tax rates. This allows us to back out the degree of foresight by equating the variance of the forecast errors from the DSGE model with the reduced-form estimates from section 2. More precisely, note that the *reduction* in the variance of the forecast error by conditioning on the risk-adjusted implicit tax rate is given by the ratio

$$\frac{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2}{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}, \{\tau_{t-j}^{RI}\}_{j=0}^{\infty}]\}^2} = \frac{\sigma_{\zeta_\tau}^2}{(\alpha_1^\tau)^2\sigma_{\zeta_\tau RI}^2 + (1-\alpha_1^\tau)^2\sigma_{\zeta_\tau}^2} = (1-\alpha_1^\tau)^{-1}. \tag{28}$$

Our definition of foresight equates conditioning on the implicit tax rate in section 2 with conditioning on the structural shocks in the DSGE models. Therefore, the mapping between the information parameter, α_1^τ, and the MA coefficient, θ, is determined by the following equality

$$\begin{aligned} \frac{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}, \{\tau_{t-j}^{RI}\}_{j=0}^{\infty}]\}^2}{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2} &= \frac{E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^{\infty}]\}^2}{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2}. \\ 1-\alpha_1^\tau &= \theta^2 \end{aligned} \tag{29}$$

Equation (29) makes clear the relationship between the reduced-form estimates of foresight given in section 2 and the calibration of the foresight intensity in the DSGE model. As the implicit tax rate becomes a perfect predictor of future tax changes, $\alpha_1^\tau \to 1$ and $\theta \to 0$, implying perfect one-step-ahead foresight, so (27) goes to zero. If there is no additional information that helps in predicting future taxes beyond the contemporaneous tax rate ($\alpha_1^\tau = 0$), then $\theta = 1$, and (23) becomes a fundamental moving average representation. Under this parameter setting, (27) shows that there is no reduction in the variance of the forecast error from observing the structural shocks $\{\varepsilon_{t-j}\}_{j=0}^{\infty}$. This is because an agent would be indifferent between observing the structural innovation ε_t and the current tax rate τ_t, since the two pieces of information are identical.

While these calculations have all been couched in the context of tax foresight, there are completely analogous representations for government spending (simply replace α_1^τ with α_1^G). Therefore, estimates of the information parameters α_1^τ and α_1^G pin down the foresight intensity—the reduction in the forecast error variance due to fiscal foresight.

This example assumes only one-period of foresight and a constant information parameter. These are unrealistic in light of the empirical evidence in section 2, which implies several quarters or years of time-varying foresight. We can nonetheless pin down the MA coefficients, θ_i^K, θ_i^L and θ_i^G for $i = 1, \ldots, q$, where q is the foresight horizon. To calibrate the high degree of tax foresight ($q = 8$), we first linearly interpolate the semi-annual estimates of α_{1t}^{τ} to get quarterly data.[17] We then use the cross-sectional average of the information parameter for municipal bond yields over one- and five-year horizons (see figure 3) during the decade of the 1980s; refer to this cross-sectional average as $\alpha_{1,t}^{\tau,1980s}$ for $t = 1980, ..., 1990$. We then use the tax events of the 1980s documented in Yang (2008) and displayed in figures 1 and 2 to back out the average (over all the tax events) of the α_1^{τ}'s that were realized one-quarter prior to implementation of the tax legislation, two-quarters prior to implementation, and so on back to five quarters prior to implementation. This procedure yields a sequence of information parameters, $\alpha_{1,T-j}^{\tau,1980s}$ where T is the date of implementation of tax event and j is the period of foresight, $j = 1, \ldots, q$.

The process for obtaining the sequence of information parameters is slightly different for government spending. Because we have the one- through five-step-ahead forecasts of real government spending from the SPF. For the high-news regime, we take the time average of α_1^G (see figure 5) from 2000 through 2009 for the one through four-step-ahead forecasts ($\alpha_{1,j}^{G,2000s}$ for $j = 1, \ldots, 4$). For the medium-news regime, we use the time average of α_1^G from 1980 through 1999 for the one through three-step-ahead forecasts ($\alpha_{1,j}^{G,1980s-1990s}$ for $j = 1, 2, 3$.).

This process generates a sequence of information parameters for both taxes and government spending where there is a unique information parameter, α_1, for each foresight horizon q. The mapping from the information parameters, the α_1's, to the moving-average coefficients, the θ's, follows the one-period example derived above, but with tedious algebra. If agents have two quarters of foresight, then the fiscal rules must have two moving-average coefficients, $\tau_t = \theta_0\varepsilon_t - \theta_1\varepsilon_{t-1} - \varepsilon_{t-2} = (L - \xi_1)(L - \xi_2)\varepsilon_t$ with $|\xi_i| < 1$ for $i = 1, 2$. The one– and two-step-ahead forecast errors must now be used to map the information parameters into the MA coefficients.

Proceeding in this fashion yields the following moving-average representations for the tax and government spending processes:[18]

Tax Foresight

- High Degree: $0.1106\varepsilon_t^i + 0.1104\varepsilon_{t-1}^i + 0.20\varepsilon_{t-2}^i + 0.1103\varepsilon_{t-3}^i + 0.1122\varepsilon_{t-4}^i + 0.1106\varepsilon_{t-5}^i + 0.109\varepsilon_{t-6}^i + 0.1115\varepsilon_{t-7}^i + 0.1141\varepsilon_{t-8}^i$
- Medium Degree: $0.208\varepsilon_t^i + 0.2042\varepsilon_{t-1}^i + 0.2011\varepsilon_{t-2}^i + 0.1961\varepsilon_{t-3}^i + 0.1912\varepsilon_{t-4}^i$
- Low Degree: $0.3324\varepsilon_t^i + 0.3333\varepsilon_{t-1}^i + 0.3342\varepsilon_{t-2}^i$

for $i \in \{L, K\}$.

[17] Our stable estimates across regimes, mean this interpolation does not effect our results. That is, we could examine a semi-annual DSGE model, as opposed to quarterly, and our conclusions would not change.

[18] We also impose the restriction that the MA coefficients sum to unity. This normalization is without loss of generality and yields the interpretation of MA coefficients as relative weights that dictate the importance of news at different horizons [see Leeper and Walker (2011)].

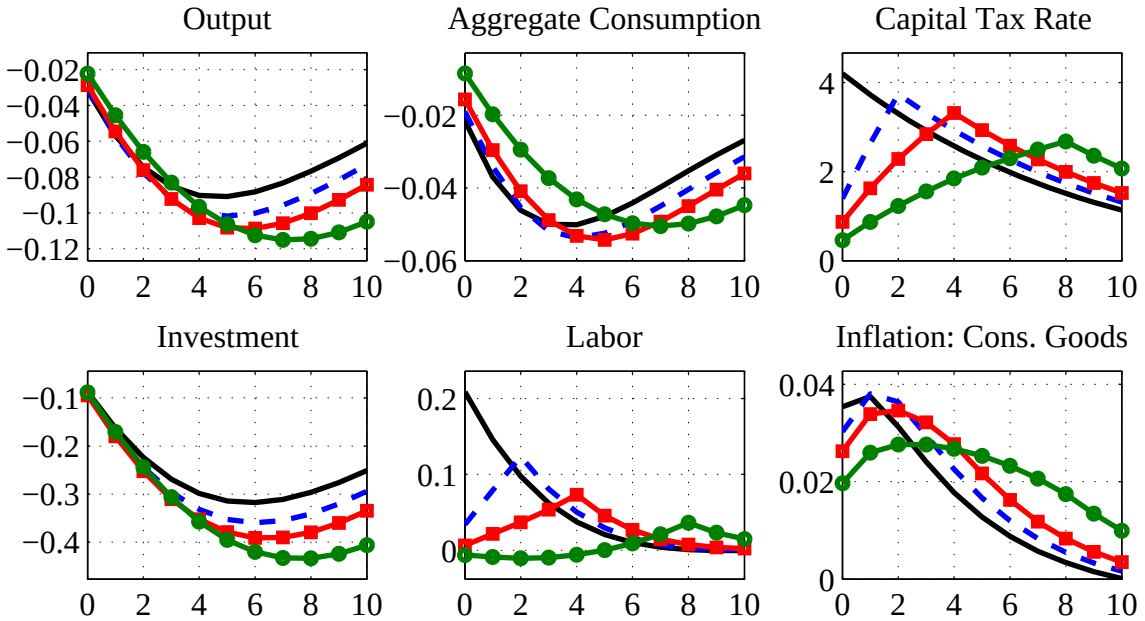


Figure 6: Response of a 1 percent increase in capital taxes. The solid line corresponds to the New Keynesian model where agents have no foresight. The other responses correspond to agents having a low degree of foresight (dashed line), a moderate degree of foresight (square markers), and a high degree of foresight (circle markers).

Government Spending

- High Degree: $0.14\varepsilon_t^G + 0.2148\varepsilon_{t-1}^G + 0.3813\varepsilon_{t-2}^G + 0.1413\varepsilon_{t-3}^G + 0.1226\varepsilon_{t-4}^G$
- Medium Degree: $0.4489\varepsilon_t^G + 0.1172\varepsilon_{t-1}^G + 0.2311\varepsilon_{t-2}^G + 0.2027\varepsilon_{t-3}^G$

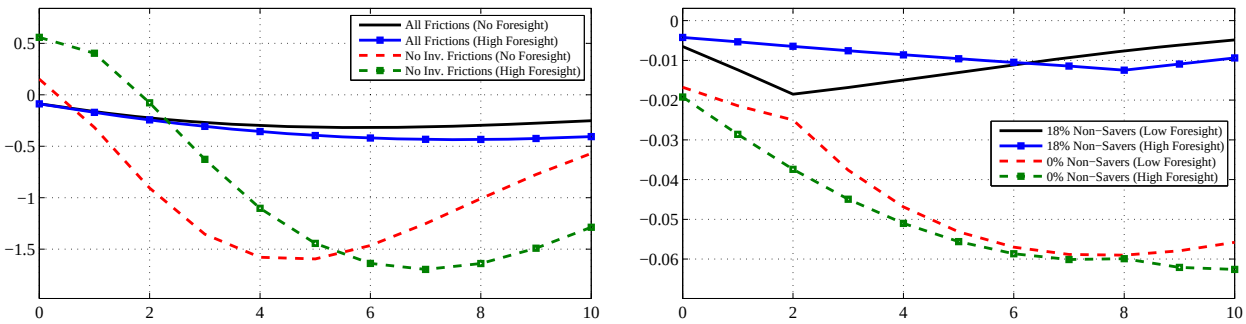
We now analyze how these different processes alter equilibrium dynamics in section 3's model.

5 Implications of Time-Varying Fiscal Foresight

This section shows how the different information regimes may lead to under-estimating the effects of news in a standard DSGE model. We use the model in section 3 and the estimated parameter values from Traum and Yang (2010) (reported in table 5, appendix B). The typical assumption that news is a time-invariant process would be innocuous if the news regimes were sufficiently close. But when the news regimes are far apart, estimating a model with time-invariant news runs of the risk of *under-reporting* the impact of fiscal foresight because the news regimes get averaged over time. The true effects of fiscal foresight may be masked by averaging high-news with no- or low-news periods. With calibrated news regimes, we can address the extent to which a standard DSGE model estimated with time-invariant news may misrepresent fiscal foresight.

5.1 Interaction with Frictions

The model includes several real frictions—investment adjustment costs, monopolistically competitive intermediate goods and labor sectors, and



(a) Investment to Capital Tax Shock (b) Employment to Labor Tax Shock

Figure 7: Figure 7a plots the response of investment to a 1 percent increase in capital taxes. The solid and square marked lines correspond to the NK model with capital utilization and investment adjustment costs with no foresight and a high degree of foresight. The dashed lines turn off the investment frictions for no foresight and high foresight. Figure 7b plots the response of employment to a 1 percent increase in labor taxes assuming a high degree of foresight. The solid line corresponds to the NK model with no non-savers. The other response assumes 18% of households are unable to save (square marker).

variable capital utilization. The motivation for many of these frictions is to smooth impulse response functions in order to better align the model with data.

Frictions smooth out the responses of agents to news about future changes to tax rates and government spending [Mertens and Ravn (2008) Schmitt-Grohé and Uribe (2008), Leeper and Walker (2011)]. Figure 6 shows the response to a capital tax shock—unanticipated and with varying degrees of foresight—in the NK model. To better understand how the frictions of the NK model interact with fiscal foresight, we plot impulse response functions with specific frictions turned off. Figure 7a plots the response of investment to a capital tax shock with investment adjustment costs and variable capital utilization turned on (solid lines) and off (dashed lines). The difference between the impulse responses for high foresight and no foresight is much larger when the frictions are turned off.

Notice that the different responses of output, investment, and aggregate consumption to different specifications of news is negligible for the first year: frictions in the NK model smooth the initial response of news shocks. At longer horizons, the differences become significant. For example, at the ten-quarter horizon, the difference between the news and no news regimes are nearly double. Firms and rational agents do not ignore the additional information provided by foresight but with adjustment costs and habit formation, the change in endogenous variables will be slow, materializing well after impact.

Conversely, because there are no frictions in the labor market, differences in labor responses show up immediately as the different news regime varies. Qualitative and quantitative differences are large for the first year, but quickly dissipate so that after 10 quarters the differences are negligible [figure 6]. Similarly, turning off these frictions, as shown by figure 7a, leads to very different responses of the news regimes at all horizons. In fact, removing the frictions can produce an investment boom on impact [Mertens and Ravn (2011)].

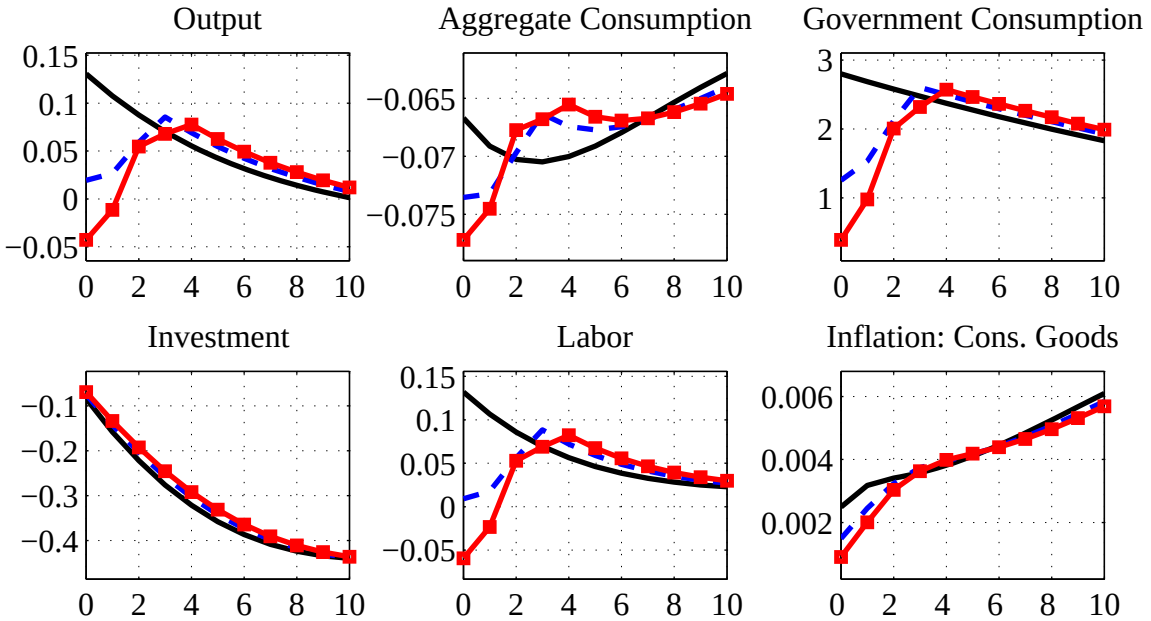


Figure 8: Response to a 1 percent increase in government consumption. The solid line corresponds to the NK model with no foresight. The dashed and square-marked lines correspond to medium and high degrees of foresight.

5.2 Percentage of Non-Savers The number of non-savers in the economy affects the extent to which time-variation in news (or news shocks in general) matters. As the proportion of non-savers increases, news shocks have less of an effect because these agents cannot take advantage of the foreknowledge of pending fiscal changes. Effects of foresight rely heavily on agents' ability to intertemporally substitute. Knowledge of a significant future increase in labor taxes has little effect on households that operate hand-to-mouth. Figure 7b shows that as a significant fraction of non-savers are added to the economy, the overall response of employment is mitigated by their inability to substitute intertemporally. This suggests that the absolute error associated with ignoring foresight, and ignoring the time-variation in foresight, is strictly decreasing in the percentage of non-savers.

Figure 7b shows that as the number of non-savers is added to the economy, the change in the impulse response of labor is a level shift towards zero. This is also the change in the impulse response as one goes from the high-news regime to the low-news regime. This suggests that the potential errors due to estimating a time-invariant news process when the actual news process is time-varying may manifest with a higher proportion of non-savers. That is, there is a potential observational equivalence between underestimating the effects of fiscal foresight and the number of non-savers. Estimating the model assuming time-invariant news will yield impulse response functions that underestimate the impact of news, which is equivalent in the model to having a high fraction of non-savers in the model.

5.3 Government Spending Figure 8 shows that government spending foresight can have large quantitative and qualitative effects in the NK model. The solid line shows the response with no foresight to an increase in government consumption. The usual result follows: invest-

ment and consumption fall as the government absorbs a larger share of goods, while output increases. With a high degree of foresight, though, output could fall in period t as agents anticipate a much higher increase in government consumption in periods $t+3$ and $t+4$. A similar result holds with respect to the response of labor; anticipated large increases in government spending in the near future cause agents to work less today. Several studies have noted that substantial foresight can lead to these qualitative differences [Mertens and Ravn (2008), Ramey (2011), Leeper et al. (2010b)].

The implication is that no-foresight fiscal multipliers are substantially different from multipliers when there is substantial foresight. DSGE environments that model the information process generating foresight explicitly potentially yield more accurate estimates of the true multiplier than, say, fiscal VARs. But this is true only if the DSGE model accurately models the news process. Ignoring time variation in news may bias the multipliers. For example, in periods of high foresight, the output multiplier at impact would be negative but averaging the three responses in figure 8 would lead to an erroneous conclusion of a positive multiplier.

6 Conclusion

By using municipal bond data and the SPF to carefully calibrate the amount of foresight in government spending and taxes, we have shown that there are periods of high news and periods of very little news. There are periods in which agents have many quarters of foresight—wars, significant changes to the tax code—and periods of little-to-no foreknowledge of pending fiscal changes. The main contribution of the paper is to show how to take reduced form estimates of news and map them into a DSGE framework. This mapping is important because we have shown within the context of a well-known DSGE model that studies that do not account for this time variation in information flows will average away the effects of news to conclude inaccurately that fiscal foresight is not relevant. Alternative news processes substantially alter equilibrium dynamics, underscoring the importance of accurately characterizing the stochastic processes governing fiscal news.

Appendix A Data Description

A.1 Municipal Bonds We utilize municipal and Treasury bond data with maturity lengths of one, five, and ten years. Yields to maturity from 1954M1 to 1994M12 on tax-exempt prime-grade general-obligation municipal bonds are obtained from Salomon Brothers' Analytical Record of Yields and Yield Spreads. Salomon Brothers' municipal data are collected on bonds of various maturity lengths on the first of each month and based on estimates of the yields of new issues sold at face value. Yields on similarly-rated (AAA) municipal bonds from 1995M1-2006M12 are obtained from Bloomberg's Municipal Fair Market Bond Index. Market yields on constant-maturity-adjusted, non-inflation-indexed U.S. Treasury securities from 1954M1-2006M12 are obtained from the Federal Reserve's Statistical Release on Selected Interest Rates. These yields reflect the average of the weekly values within each month, which are interpolated from the daily yield curve.

A.2 Government Spending Data on quarterly nominal federal government consumption and gross investment spending from 1981Q3 to 2010Q1 are obtained from the National Income and Product Accounts, published by the Bureau of Economic Analysis (BEA). A real series of federal government consumption and gross investment expenditures in chained 2005 dollars (RGFED) was generated using the component-specific real GDP quantity index (QI) [NIPA Table 1.1.3, line 22] and annual component-specific nominal GDP (NGFED) [NIPA Table 1.1.5, line 22]. The following formula was applied to convert from current dollars to chained 2005 dollars:

$$\mathrm{RGFED}^{Q}_{\mathrm{BY}} = \left(\frac{\mathrm{QI}^{Q}_{\mathrm{CY}}}{\mathrm{QI}^{A}_{\mathrm{BY}}}\right)\mathrm{NGFED}^{Q}_{\mathrm{BY}},$$

where A and Q designate between annual and quarterly values and CY and BY denote current quarterly and base year (annual) values.

A.3 Survey of Professional Forecasters Mean *forecasts* of real federal government consumption and gross investment from 1981Q3 to 2010Q1 over one, two, three, four, and five year horizons are taken from the Survey of Professional Forecasters (SPF), conducted by the Federal Reserve Bank of Philadelphia. Unfortunately, the published data is not provided under a constant base year and is affected by several changes in the base year set by the BEA. This creates two complications. First, the BEA does not publish price indexes corresponding to historical base years. Second, the components of and the methodology for collecting federal government spending data has changed over time. In the first quarter of 1996, the BEA's price and quantity indexes switched to chain-weighted measures. Moreover, in the same quarter, government purchases were replaced by government consumption and gross investment spending, which lead to a substantial upward revision in the government component of GDP.[19] These changes forced us to employ two different methods to transform this series of forecasts into constant 2005 dollars.

Between 1981Q3 and 1995Q4, we collect nominal government purchases (Table 1) and the component-specific implicit price deflator (Table 7.1) from quarterly issues of the Survey

[19] For more details surrounding the precise changes in the definition of government spending see the Survey of Business issues from September 1995 and January 1996.

Table 4: Base Years for NIPA Variables in the SPF

Range of Surveys	Base Year
1996Q1 to 1999Q3	1992
1999Q4 to 2003Q4	1996
2004Q1 to 2009Q2	2000
2009Q3 to 2010Q1	2005

of Current Business, which were downloaded from the Federal Reserve Archival System for Economic Research. A time series of these variables was created using the most recently revised estimates. Real forecasts were then converted to current dollars by multiplying the quarterly real forecast by the quarterly implicit price deflator and dividing by 100. To account for the change in the definition of government spending, we collect current data on nominal federal government consumption and gross investment and calculate the difference from the past definition. We then scale up the calculated nominal forecasts to obtain government spending forecasts based on its new definition. Finally to convert these values into constant 2005 dollars, we multiply by 100 and divide the corresponding quarterly implicit price deflator.

Between 1996Q1 and 2010Q1, the data is first converted to current dollars by constructing the component-specific implicit price deflator (IPD) in each of the relevant base years. To re-base the index, we applied the following transformation

$$\text{NIPD}^{\text{Q}}_{\text{CY}} = \frac{\text{OIPD}^{\text{A}}_{\text{CY}}}{\text{OIPD}^{A}_{\text{NBY}}},$$

where NIPD and OIPD correspond to the implicit price deflator series under the new and old base years and NBY stands for the new (desired) base year. We then construct a new IPD series with base years corresponding to the data specified in table 4. Using the generated series, we obtain nominal forecasts by multiplying each quarterly data point by the current implicit price deflator with the appropriate base year. The constructed nominal series is then converted to constant 2005 dollars using the same procedure that was applied to pre-1996 data.

A.4 Marginal Tax Rates

Marginal income tax rates for married individuals filing joint returns are obtained from Internal Revenue Service publications and the Tax Policy Center. Following Fortune (1996), marginal tax brackets, reported in current dollars, are converted to constant 1980 dollars using the implicit price deflator [NIPA Table 1.1.9]. A series of actual and ex post tax rates are then constructed using marginal tax rates for investors earning \$100,000, \$75,000, and \$50,000 annually in constant 1980 dollars. Annual tax rates are then applied to each month of each corresponding year.

Appendix B Parameter Values

This appendix reports the parameter values estimated by Traum and Yang (2010)

Table 5: New Keynesian Model Parameters

Baseline Calibration		
Quarterly discount factor	β	0.99
Capital share	α	0.36
Private capital depreciation rate	δ_0	0.025
Government capital depreciation rate	δ^G	0.02
Elasticity of substitution between labor inputs	ψ^w	8
Elasticity of substitution between intermediate goods	ψ^p	8
Steady-state inflation rate	π	1
Government consumption-to-output ratio	s^{GC}	0.0698
Government investment-to-output ratio	s^{GI}	0.00395
Quarterly debt-to-output ratio	s^b	1.5436
Steady-state labor tax rate	τ^L	0.209
Steady-state capital tax rate	τ^K	0.196
Steady-state consumption tax rate	τ^C	0.0149
Elasticity of output with respect to public capital	α^G	0.05
Implied Parameters		
Capital-to-output ratio	K/Y	7.2152
Consumption-to-output ratio	C/Y	0.7459
Labor-to-output ratio	L/Y	0.3547
Nominal wage rate	w	1.557
Transfers-to-output ratio	s^{TR}	0.0978
Savers consumption-to-output ratio	C^S/Y	0.7923
Non-savers consumption-to-output ratio	C^N/Y	0.5342
Estimated Parameters		
Risk aversion	γ	2.7
Inverse Frisch elasticity of labor	κ	2.1
Fraction of savers	μ	0.18
Wage stickiness	ω_w	0.69
Price stickiness	ω_p	0.82
Capital utilization	ψ	0.38
Investment adjustment costs	s	7.4
Wage partial indexation	χ^w	0.39
Price partial indexation	χ^p	0.31
Government consumption response to debt	γ_{GC}	0.17
Government investment response to debt	γ_{GI}	0.0033
Capital tax response to debt	γ_K	0.17
Labor tax response to debt	γ_L	0.16
Transfers response to debt	γ_Z	0.074
Capital response to output	φ_K	0.78
Labor response to output	φ_L	0.43
Interest rate response to inflation	ϕ_π	1.9
Interest rate response to output	ϕ_y	0.095
Lagged interest rate response	ρ_r	0.86
Persistence of government consumption shock	ρ_{GC}	0.96
Persistence of government investment shock	ρ_{GI}	0.76
Persistence of capital tax shock	ρ_K	0.89
Persistence of labor tax shock	ρ_L	0.94
Persistence of consumption tax shock	ρ_C	0.90
Persistence of transfers shock	ρ_Z	0.79
Std. Dev of government consumption shock	σ_{GC}	2.8
Std. Dev of government investment shock	σ_{GI}	4
Std. Dev of capital tax shock	σ_K	4.2
Std. Dev of labor tax shock	σ_L	2.3
Std. Dev of consumption shock	σ_C	3.3
Std. Dev of transfers shock	σ_Z	2.6
Co-movement between capital and labor taxes	ϕ_{KL}	0.23

References

Ang, A., V. Bhansali, and Y. Xing (2010): "Taxes on tax-exempt bonds," *The Journal of Finance*, 65, 565–601.

Auerbach, A. J. and J. Slemrod (1997): "The Economic Effects of the Tax Reform Act of 1986," *Journal of Economic Literature*, 35, 589–632.

Barsky, R. B. and E. R. Sims (2011): "News Shocks and Business Cycles," *Journal of Monetary Economics*, forthcoming.

Beaudry, P. and F. Portier (2006): "Stock Prices, News, and Economic Fluctuations," *American Economic Review*, 96, 1293–1307.

Branson, W. H., A. Fraga, and R. A. Johnson (1986): "Expected Fiscal Policy and the Recession of 1982," in *Prices, Competition, and Equilibrium*, ed. by M. H. Peston and R. E. Quandt, Totowa, New Jersey: Barnes & Noble, 109–128.

Chalmers, J. M. R. (1998): "Default Risk Cannot Explain the Muni Puzzle: Evidence from Municipal Bonds That Are Secured by U.S. Treasury Obligations," *Review of Financial Studies*, 11, 281–308.

Christiano, L. J., C. Ilut, R. Motto, and M. Rostagno (2008): "Monetary Policy and Stock Market Boom-Bust Cycles," Working Paper No. 955, European Central Bank.

Comin, D., M. Gertler, and A. M. Santacreu (2008): "News, Technology Adoption and Economic Fluctuations," Manuscript, New York University.

Durbin, J., A. Harvey, S. Koopman, and N. Shephard (2004): *State Space and Unobserved Component Models: Theory and Applications*, Cambridge University Press.

Fernández-Villaverde, J., J. F. Rubio-Ramírez, T. J. Sargent, and M. W. Watson (2007): "ABCs (and Ds) of Understanding VARs," *American Economic Review*, 97, 1021–1026.

Fisher, J. D. and R. Peters (2010): "Using Stock Returns to Identify Government Spending Shocks," *Economic Journal*, 120, 414–436.

Fortune, P. (1996): "Do Municipal Bond Yields Forecast Tax Policy?" *New England Economic Review*, 29–48.

Fujiwara, I., Y. Hirose, and M. Shintani (2008): "Can News Be a Major Source of Aggregate Fluctuations? A Bayesian DSGE Approach," Bank of Japan Discussion Paper Series 2008-E-16.

Hall, R. E. (1971): "The Dynamic Effects of Fiscal Policy in an Economy with Foresight," *Review of Economic Studies*, 38, 229–244.

Hansen, L. P. and T. J. Sargent (1980): "Formulating and Estimating Dynamic Linear Rational Expectations Models," *Journal of Economic Dynamics and Control*, 2, 7–46.

——— (1991): "Two Difficulties in Interpreting Vector Autoregressions," in *Rational Expectations Econometrics*, ed. by L. P. Hansen and T. J. Sargent, Boulder, CO: Westview Press, 77–119.

Jaimovich, N. and S. Rebelo (2009): "Can News About the Future Drive the Business Cycle?" *American Economic Review*, 99, 1097–1118.

Judd, K. L. (1985): "Short-Run Analysis of Fiscal Policy in a Simple Perfect Foresight Model," *Journal of Political Economy*, 93, 298–319.

Kochin, L. A. and R. W. Parks (1988): "Was the Tax-Exempt Bond Market Inefficent of Were Future Expected Tax Rates Negative," *The Journal of Finance*, 43, 913–931.

Kriwoluzky, A. (2009): "Pre-Announcement and Timing—The Effects of a Goverment Expenditure Shock," EUI Working Paper MWP 2009/40.

Leeper, E. and T. Walker (2011): "Information Flows and News Driven Business Cycles," *Review of Economic Dynamics*, 14, 55–71.

Leeper, E. M. (1989): "Policy Rules, Information, and Fiscal Effects in a 'Ricardian' Model," Federal Reserve Board, International Finance Discussion Paper No. 360.

Leeper, E. M., M. Plante, and N. Traum (2010a): "Dynamics of fiscal financing in the United States," *Journal of Econometrics*, 156, 304–321.

Leeper, E. M., T. B. Walker, and S.-C. S. Yang (2008): "Fiscal Foresight: Analytics and Econometrics," NBER Working Paper No. 14028.

——— (2010b): "Government investment and fiscal stimulus," *Journal of Monetary Economics*, 57, 1000–1012.

——— (2011): "Foresight and Information Flows," NBER Working Paper No. 16951.

Lippi, M. and L. Reichlin (1994): "VAR Analysis, Nonfundamental Representations, Blaschke Matrices," *Journal of Econometrics*, 63, 307–325.

Mertens, K. and M. Ravn (2008): "The Aggregate Effects of Anticipated and Unanticipated U.S. Tax Policy Shocks: Theory and Empirical Evidence," EUI Working Papers, ECO 2008/05.

——— (2010): "Measuring the Impact of Fiscal Policy in the Face of Anticipation: A Structural VAR Approach," *Economic Journal*, 120, 393–413.

Mertens, K. and M. O. Ravn (2011): "Understanding the Aggregate Effects of Anticipated and Unanticipated Tax Policy Shocks," *Review of Economic Dynamics*, 14, 27–54.

Park, S. (1997): "The Relationship Between Government Financial Condition and Expected Tax Rates Reflected in Municipal Bond Yields," *National Tax Journal*, 50, 23–38.

Perotti, R. (2011): "Expectations and Fiscal Policy: An Empirical Investigation," Manuscript, Università Bocconi.

Poterba, J. M. (1986): "Explaining the Yield Spread between Taxable and Tax-Exempt Bonds: The Role of Expected Tax Policy," in *Studies in State and Local Public Finance*, Chicago, IL: University of Chicago Press, 5–51.

——— (1988): "Are Consumer Forward Looking? Evidence from Fiscal Experiments," *American Economic Review Papers and Proceedings*, 78, 413–418.

——— (1989): "Tax Reform and the Market for Tax-Exempt Debt," *Regional Science and Urban Economics*, 19, 537–562.

Ramey, V. A. (2009): "Defense News Shocks, 1939-2008: An Analysis Based on News Sources," Manuscript, University of California, San Diego.

——— (2011): "Identifying Government Spending Shocks: It's All in the Timing," *Quarterly Journal of Economics*, 126, 1–50.

Schmitt-Grohé, S. and M. Uribe (2008): "What's 'News' in Business Cycles?" NBER Working Paper No. 14215.

Sims, C. A. (1988): "Identifying Policy Effects," in *Empirical Macroeconomics for Interdependent Economics*, ed. by R. C. Bryant, Washington, DC: The Brookings Institution, 308–321.

Traum, N. and S.-C. S. Yang (2010): "When Does Government Debt Crowd Out Investment?" CAEPR Working Paper No. 2010-006.

Yang, S.-C. S. (2005): "Quantifying Tax Effects Under Policy Foresight," *Journal of Monetary Economics*, 52, 1557–1568.

——— (2008): "A Chronology of Postwar U.S. Federal Income Tax Policy," CAEPR Working Paper No. 2007-021.