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 and 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 argue that news about fiscal policy is a time-varying process. We demonstrate that the ignoring the time variation can have important consequences in a conventional macroeconomic model.
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, 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 and 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 antici-
pated change 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. We then derive a unique mapping from these estimates of fiscal foresight to a standard DSGE model (Section 4). 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.

We argue that 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 medium-to-low or no foresight. This result implies that models that do not account for the time-varying nature of information flows will average away the effects of news. These studies might incorrectly conclude that fiscal foresight is not relevant for explaining business cycle dynamics and would be unable to assess the effects of fiscal 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. However, we argue that the typical assumption of news as a time-invariant process runs the risk of under-reporting the impact of fiscal foresight due to the time-averaging of news periods. That is, the true effects of fiscal foresight may be masked by averaging high news periods with periods of no or low news. Given that we have on hand a calibration of news regimes, we address the extent to which a standard DSGE model estimated with time-invariant news may misrepresent fiscal foresight.

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. However when estimating a DSGE model using likelihood or Bayesian methods, 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. 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

\footnote{One can even use this information to back out the true structural innovations from a VAR [Mertens and Ravn (2010)].}
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, 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. The differential treatment of municipal and treasury bonds has useful implications for identifying news about tax changes. If $Y^M_t$ is the yield on a municipal bond at $t$ and $Y_t$ 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 $\tau^I_t = 1 - Y^M_t / Y_t$. This is the tax rate at which the investor is indifferent between the 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)]. 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. While over half of the estimated coefficients have the correct sign and are statistically significant for bonds with a 1-year maturity, only five of the 22 tax events meet these criteria at 10-year horizons. Evidently, municipal bond yield yields are more informative about tax events in the near future 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, and so forth) grow with the horizon being considered. Although event studies

\footnote{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.}

\footnote{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

<table>
<thead>
<tr>
<th>Predicted Sign</th>
<th>Estimation Results</th>
<th>Date</th>
<th>Event</th>
</tr>
</thead>
<tbody>
<tr>
<td>Positive</td>
<td>c,s</td>
<td>c</td>
<td>March 1954</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>c</td>
<td>Sept. 1963</td>
</tr>
<tr>
<td>Positive</td>
<td>c,s</td>
<td>c</td>
<td>Feb. 1968</td>
</tr>
<tr>
<td>Positive</td>
<td>c,s</td>
<td>x</td>
<td>May 1969</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>c</td>
<td>Aug. 1969</td>
</tr>
<tr>
<td>Negative</td>
<td>c</td>
<td>x</td>
<td>Sept. 1971</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>x</td>
<td>Jan. 1975</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>c</td>
<td>Oct. 1975</td>
</tr>
<tr>
<td>Negative</td>
<td>x,s</td>
<td>x</td>
<td>Aug. 1976</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>c</td>
<td>March 1977</td>
</tr>
<tr>
<td>Negative</td>
<td>x</td>
<td>x</td>
<td>Aug. 1978</td>
</tr>
<tr>
<td>Negative</td>
<td>x</td>
<td>c</td>
<td>July 1981</td>
</tr>
<tr>
<td>Negative</td>
<td>c,s</td>
<td>c,s</td>
<td>July 1982</td>
</tr>
<tr>
<td>Positive</td>
<td>c,s</td>
<td>c,s</td>
<td>March 1984</td>
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<tr>
<td>Negative</td>
<td>c,s</td>
<td>c,s</td>
<td>Dec. 1985</td>
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<tr>
<td>Negative</td>
<td>x,s</td>
<td>c,s</td>
<td>June 1986</td>
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<td>Negative</td>
<td>c</td>
<td>c,s</td>
<td>Sept. 1990</td>
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<tr>
<td>Positive</td>
<td>c,s</td>
<td>c</td>
<td>May 1993</td>
</tr>
<tr>
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<td>c</td>
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<td>June 1997</td>
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<tr>
<td>Negative</td>
<td>x</td>
<td>x,s</td>
<td>May. 2001</td>
</tr>
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<td>x,s</td>
<td>x</td>
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<tr>
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<td>x,s</td>
<td>c,s</td>
<td>May 2004</td>
</tr>
<tr>
<td>Negative</td>
<td>x,s</td>
<td>c,s</td>
<td>July 2004</td>
</tr>
</tbody>
</table>

* 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.

suffer from the drawback of requiring the 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

\[
1 = \sum_{t=1}^{T} \frac{C^M}{(1 + R^T_t)^t} + \frac{1}{(1 + R^T_T)^T},
\]

where $R^T_t$ 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

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

(2)

where $\tau_t^e$ 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

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

(3)

where $\omega_t = \delta_t/\sum_{j=1}^{T} \delta_j$ and $\delta_t = (1 + R_t^e)^{-t}$. Because the $\omega$ 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\(^5\)

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

(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. In principle, 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.

\(^5\)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 $\sum_{j=s+1}^{t} \delta_j \tau_j^e = \tau_t^I \sum_{j=1}^{s} \delta_j - \tau_s^I \sum_{j=1}^{s} \delta_j$. Dividing 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.
By the time the tax reform actually took effect, agents had factored the entire effect of the policy 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. To account for all unobservable risk that may arise, Fortune (1996) introduces a time-invariant “quality premium”, $\xi$, in the relationship between yields on municipal bonds and treasuries. The risk-adjusted implicit tax rate is given by

$$\tau_{t}^{RI} = 1 - \frac{Y_{t}^{M} + \xi}{Y_{t}^{t}}. \tag{5}$$

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 $\tau_{t+i}$ 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, $\tau_{t}^{RI}$, and the spot tax rate, $\tau_{t}$, along with a forecast error to obtain

$$T_{t} = \alpha_{1}^{\tau_{t}} \tau_{t}^{RI} + (1 - \alpha_{1}^{\tau_{t}}) \tau_{t} + \varepsilon_{t}^{\tau}, \quad \varepsilon_{t}^{\tau} \sim N(0, \sigma_{\tau}^{2}). \tag{6}$$

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 ($\zeta_{\tau,RI} = \ldots$

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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.
Let $\zeta_T$ denote the forecast error from predicting changes in the ex-post tax rate, conditional on the spot tax rate alone ($\zeta_T = T - \tau$). The composite forecast error is given by the convex combination of the two, $\varepsilon_T = \alpha_T^1 \zeta_T + (1 - \alpha_T^1) \zeta_T$. The optimal weight, $\alpha_T^1$, is chosen to minimize the variance of the forecast error. This weight is given by

$$\alpha_T^1 = \frac{\sigma_{\zeta_T}^2}{\sigma_{\zeta_T}^2 + \sigma_{\zeta_T, RI}^2}, \quad (7)$$

where $\sigma_{\zeta_T}^2$ and $\sigma_{\zeta_T, RI}^2$ are the variances of the forecast errors $\zeta_T$ and $\zeta_T, RI$. 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, $\sigma_{\zeta_T, RI}^2$, would be zero and $\alpha_T^1$ would be set to unity.

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

$$T_t - \tau_t = \alpha_T^1 (\tau_I^t - \tau_t) + \alpha_T^2 (1/Y_t) + \varepsilon_T^t, \quad (8)$$

where $\alpha_T^1$ measures the information content of municipal bonds and $\alpha_T^2 = -\alpha_T^1 \xi$ 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

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\(^9\)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.
information parameter, $\alpha_1$, 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, $\alpha_2$, 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

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

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.\footnote{See Durbin et al. (2004) for more details on the estimation procedure.}

Table 3 reports the estimation allowing for time-varying parameter values. Notice that the standard deviation is largest for the information parameter ($\delta_{1,t}$). 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 five quarters.\footnote{For in-depth analysis on the explanatory power of the SPF see Perotti (2011).}
Figure 3: Estimated path of the time-varying information parameter $\alpha^*_T$. 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.

Figure 4: Time-varying information parameter, $\alpha^*_T$, for 1, 2, 3, and 4-step-ahead forecasts of real government spending from 1982-2009.
nominal federal government consumption and gross investment spending over the same period are obtained from the National Income and Product Accounts, published by the Bureau of Economic Analysis (BEA). A real series of 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 5 plots real federal government spending along with Ramey’s war dates. As is evident from this picture, defense news is predominately followed by stark changes in government consumption and investment expenditures.

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

\[ G_{t+j} = \alpha_t^G G_{t+j|t} + (1 - \alpha_t^G) \rho_t^G G_t + \varepsilon_t^G, \quad j = 1, \ldots, 5, \quad \varepsilon_t^G \sim N(0, \sigma_G^2) \]  
\[ \alpha_t^G = \alpha_{t-1}^G + \eta_t^G, \quad \eta_t^G \sim N(0, \delta_G^2). \]  

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. We fit the AR(1) model using OLS on quarterly first-differenced real federal government expenditures from 1981Q3 to 2010Q1.\textsuperscript{12} Analogous to the tax foresight case and (7), \( \alpha^G \) is determined by whichever forecast has the smaller forecast error variance.

As with tax foresight, we allow the information parameter for government spending to be a time-varying process. Once again, we use a Kalman filter to back out the path of the information parameter. Figure 4 plots the \( \alpha_t^G \) parameter for \( j = 1, 2, 3, 4 \) from 1982

\textsuperscript{12}More elaborate time series specifications for the government spending process were estimated but model selection criteria (e.g., AIC) preferred the AR(1) specification.
through 2009.\textsuperscript{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 5.\textsuperscript{14}

3 Model for Policy Evaluation

We adopt a conventional new Keynesian model based on Traum and Yang (2010) that
incorporates several frictions that have become standard in the literature.\textsuperscript{15}

The model includes two types of households: savers, denoted by $S$, who have access to a
complete set of contingent claims, and non-savers, denoted by $N$, who each period consume
their entire disposable income. A fraction $\mu \in [0, 1]$ of the population is savers and the
remaining $1 - \mu$ fraction is non-savers. The continuum of agents have common preferences,
as represented by those of agent $j \in [0, 1]$

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

(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 monop-
olistically competitive setting. The total amount of labor supplied by household $j$ satisfies
$L_t^A(j) = \int_0^1 \ell_t^A(j, \ell) d\ell$, where $\ell_t^A(j, \ell)$ is the amount of labor of input type $\ell$ supplied by agent
$j$ of type $A$.

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

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

(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 $\tau_t^L$, $\tau_t^K$, and $\tau_t^C$ 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

$$
H_t(j) \equiv (1 - \tau_t^L) \int_0^1 \frac{W_t(\ell)}{P_t} \ell_t^S(j, \ell) d\ell + z_t(j) + d_t(j),
$$

(14)

\textsuperscript{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.

\textsuperscript{14}Our estimates of $\alpha_t^G$ at a 5-quarter horizon showed little informational content and are, therefore, omitted
from Figure 4.

\textsuperscript{15}To simplify the presentation of the model, we omit several of the shocks that Traum and Yang needed
for estimation purposes.
where $W_t(\ell)$ is the nominal wage for labor type $\ell$, $z_t(j)$ 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

$$k_t(j) = (1 - \delta[v_t(j)])k_{t-1}(j) + \left[ 1 - s \left( \frac{i_t(j)}{k_{t-1}(j)} \right) \right] i_t(j), \quad (15)$$

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, $\delta$, is positively related to the utilization rate, $v_t$, and is given by

$$\delta[v_t(j)] = \delta_0 + \delta_1(v_t(j) - 1) + \frac{\delta_2}{2}(v_t(j) - 1)^2, \quad (16)$$

where $\delta_0$, $\delta_1$, and $\delta_2$ are calibrated parameters.\(^{16}\)

The budget constraint for non-saver $j \in (1 - \mu, 1]$, who does not have access to asset markets, is

$$c_t^N(j) = (1 - \tau_t^L) \int_0^1 \frac{W_t(\ell)}{P_t} \ell_t^N(j, \ell) d\ell + z_t(j). \quad (17)$$

Aggregate demand for labor services is not biased toward a certain labor type, $A$. Therefore, in equilibrium 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, $L_t = \left[ \int_0^1 \ell_t(\ell)^{1/(1+\eta^w)} d\ell \right]^{1+\eta^w}$, where $\eta^w$ 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 $l_t(\ell) = L_t^d(W_t(\ell)/W_t)^{(1+\eta^w)/\eta^w}$, 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

$$y_t(i) = (v_t k_{t-1}(i))^\alpha (\ell_t(i))^{1-\alpha} (K_{t-1}^G)^{\alpha^G}, \quad (18)$$

where $k_t(i)$ and $\ell_t(i)$ denote the capital stock and level of employment used by firm $i$, $\alpha \in [0, 1]$ is the cost share of capital, and $\alpha^G$ is the elasticity of output with respect to the stock of government capital $K_{t-1}^G$.

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

\(^{16}\delta_1\) is calibrated so that $v = 1$ in steady-state. The parameter $\psi \in [0, 1]$ is defined so that $\delta''(1)/\delta'(1) = \delta_2/\delta_1 \equiv \psi/(1 - \psi)$.\[13]
the price of the final good, and \( \psi \equiv (1 + \eta^p)/\eta^p \) is the elasticity of substitution between intermediate goods.

Both wages and prices adjust according to a Calvo pricing mechanism. Each period, a union has the opportunity to adjust the nominal wage rate with probability \((1 - \omega^w)\). In the event that the union does not receive a pricing signal, wages are indexed to inflation according to the rule

\[
W_t(l) = W_{t-1}(l)\pi_{t-1}^w,
\]

where \( \chi^w \) parameterizes the degree of wage indexation. If, on the other hand, a union is fortunate enough to be able to freely adjust the nominal wage rate, \( \tilde{W}_t(l) \), to maximize the lifetime utility of households given by

\[
E_t \sum_{i=0}^{\infty} (\beta \omega_w)^i \left[ (1 - \mu) \frac{(\tau^S_{t+1})^{1-\gamma} - 1}{1 - \gamma} + \mu \frac{(\tau^N_{t+1})^{1-\gamma} - 1}{1 - \gamma} - \frac{L_{t+1}^{1+\kappa}}{1 + \kappa} \right],
\]

subject to the aggregate budget constraints for both savers and non-savers and the individual and aggregate labor demand functions. In a symmetric equilibrium, where \( \tilde{W}_t(l) = \tilde{W}_t \), the aggregate nominal wage is

\[
W_t = \left[ (1 - \omega^w)\tilde{W}^{-\frac{1}{\eta^p}} + \omega_w(\pi_{t-1}^w)^{-\frac{\chi^w}{\eta^w}} W_{t-1}^{-\frac{1}{\eta^w}} \right]^{-\eta^w},
\]

where \( \pi_{t-1}^w = W_t/W_{t-1} \) is the gross wage inflation rate.

Similarly, each intermediate goods producing firm may reset its price only with probability \((1 - \omega^p)\) in any given period. Firms that are unable to make optimal adjustments simply index their price level to past inflation by setting

\[
\bar{p}_t(i) = \bar{p}_{t-1}(i)\pi_{t-1}^p,
\]

where \( \chi^w \) parameterizes the degree of price indexation. 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. In a symmetric equilibrium, where \( \bar{p}_t(i) = \bar{p}_t \), the producer price index, \( \bar{P}_t \), evolves according to

\[
\bar{P}_t = \left[ (1 - \omega^p)\bar{p}_t^{-\frac{1}{\eta^p}} + \omega_p\pi_{t-1}^{\chi^p} \bar{P}_{t-1}^{-\frac{1}{\eta^p}} \right]^{-\eta^p}.
\]

The fiscal authority finances government consumption, \( G_t \), government investment, \( G^I_t \), and government transfers, \( Z_t \), through proportional taxes levied against consumption, labor income, and capital returns and by issuing one-period nominal debt. The government’s flow budget constraint is

\[
B_t + \tau^K_t \frac{R^K_t}{\bar{P}_t} v_t K_{t-1} + \tau^L_t \frac{W_t}{\bar{P}_t} L_t + \frac{\tau^C_t}{1 + \tau^C_t} C_t = \frac{R_{t-1} B_{t-1}}{\pi_t} + G_t + G^I_t + Z_t.
\]

Productive government capital evolves according to

\[
K^G_t = (1 - \delta^G) K^G_{t-1} + G^I_t.
\]
Fiscal variables are determined by the following rules

\begin{align}
\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^{b} \right) + \phi_{KL} \sigma_L \varepsilon_t^L + \sum_{i=0}^{q} \theta_i^K \varepsilon_{t-i} \tag{26} \\
\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^{b} \right) + \phi_{KL} \sigma_K \varepsilon_t^K + \sum_{i=0}^{q} \theta_i^L \varepsilon_{t-i} \tag{27} \\
\hat{\gamma}_t &= \rho_G \hat{\gamma}_{t-1} - (1 - \rho_G) \gamma_G \hat{s}_t^{b} + \sum_{i=0}^{q} \theta_i^G \varepsilon_{t-i} \tag{28} \\
\hat{\gamma}_t^I &= \rho_I \hat{\gamma}_{t-1}^I - (1 - \rho_I) \gamma_I \hat{s}_t^{b} + \sigma_I \varepsilon_t^I \tag{29} \\
\hat{Z}_t &= \rho_Z \hat{Z}_{t-1} - (1 - \rho_Z) \gamma_Z \hat{s}_t^{b} + \sigma_Z \varepsilon_t^Z \tag{30} \\
\hat{\tau}_t^C &= \rho_C \hat{\tau}_{t-1}^C + \sigma_C \varepsilon_t^C \tag{31}
\end{align}

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, \( s_{t-1}^b \), 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 \( \sum_{i=0}^{q} \theta_i^j \varepsilon_{t-i} \) for \( j \in \{K, L, G\} \), reflect fiscal foresight. Finally, consistent with actual U.S. tax behavior, capital and labor taxes are permitted to be correlated through the parameter \( \phi_{KL} \).

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

\begin{align}
\hat{R}_t &= \rho_r \hat{R}_{t-1} + (1 - \rho_r) \left[ \varphi_r \hat{\pi}_t + \phi_y \hat{\pi}_t \right], \tag{32}
\end{align}

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

For a generic type-specific variable \( x_t \), its aggregate counterpart is given by \( X_t \). Aggregate consumption, which is composed of consumption by both savers and non-savers is given by

\begin{align}
C_t = \int_0^1 c_t(j) dj = (1 - \mu) c_t^S + \mu c_t^N.
\end{align}

Lump-sum transfers are identical across households so that

\begin{align}
Z_t = \int_0^1 z_t(j) dj = z_t.
\end{align}

Since non-savers do not have access to asset markets, the aggregate levels of bonds, investment, capital, and dividends are given by

\begin{align}
B_t &= \int_0^1 b_t(j) dj = (1 - \mu) b_t, \quad K_t = \int_0^1 k_t(j) dj = (1 - \mu) k_t, \\
I_t &= \int_0^1 i_t(j) dj = (1 - \mu) i_t, \quad D_t = \int_0^1 d_t(j) dj = (1 - \mu) d_t.
\end{align}

To close the model, the aggregate resource constraint is

\begin{align}
Y_t = C_t + I_t + G_t + G_t^I. \tag{33}
\end{align}
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. One of the main contributions of this paper is to extract the amount of foresight from the reduced-form estimates of Section 2, and to map these estimates into the New Keynesian model of 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.

As an example of foresight horizon, consider the Economic Recovery Tax Act of 1981. The foresight horizon can be separated into two periods: between initial proposal and final enactment—or rejection—of a new tax law (“inside lag”) and between enactment and when the law takes effect (“outside lag”). During the inside lag, agents are forming expectations about the likelihood and precise form of proposed legislation. In announcing his candidacy for president in November 1979, Ronald Reagan made clear that he intended to substantially lower taxes: “The key to restoring the health of the economy lies in cutting taxes” [Reagan (1979)]. News about future taxes, then, arrived throughout 1980, evolving with Reagan’s prospects for winning office. An additional six months passed between President Reagan’s formal call for tax relief in February 1981 and the legislation’s enactment. The inside lag associated with this tax change is, arguably, five or more quarters. The Economic Recovery Tax Act of 1981, enacted in August 1981, phased in tax reductions through the beginning of 1984 to yield an outside lag of 10 quarters. In the calibration described more thoroughly below, we use Yang’s 2008 chronology to calibrate the foresight horizon. Yang proxies the inside and outside lags by defining the total lag as the number of quarters elapsed between when a tax proposal is first announced by a President and when it takes effect. This is a conservative estimate of the foresight horizon as it does not take into consideration tax proposals announced by candidates for president.

An example of foresight intensity pertaining to government spending is the recently enacted American Recovery and Reinvestment Act (ARRA) of 2009. Prior to the passage of the bill, agents did not know the size or composition of the anticipated government spending. Table 4 contains the Congressional Budget Office’s estimates of costs and outlays associated with two pieces of legislation involving government investment. Based on historical spending rates, the CBO assumes that outlays for government investment take place over several years following the authorization. For the ARRA, Congress authorized $27.5 billion for highway construction in 2009, yet the estimated outlays are only $2.75 billion for fiscal year 2009. Another example is the National Highway Bridge Reconstruction and Inspection Act (NHBRIA) of 2008, which was not enacted but would have authorized appropriations of about $1 billion in fiscal year 2009 for repairing, rehabilitating, and replacing bridges on

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17These labels date back to Friedman (1948), where we combine the “recognition” and “decision” lags to form inside lags and our outside lags refer to how long it takes legislation to change tax rates.
public roadways. Outlays associated with this legislation were planned to extend more than four years into the future. Due to the differences between outlays and authorized spending, agents have a precise measure of the projected increase in government spending that can be attributed to the ARRA over the next several years.

### 4.1 Foresight Horizon

By allowing for time variation in the information parameters, $\alpha_{1T}$ and $\alpha_{1G}$, we are able to 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 4 that there is substantial time variation in the news content of municipal bonds and the SPF, and the high news (high $\alpha_{1}$) 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. More specifically, we calibrate the high degree of tax foresight specification using data from the 1980s. The 1980s was 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 time (under one year on average). Interestingly, as evidenced by Figure 3, 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...
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 4, the information content of the SPF’s forecasts of changes in government spending over 1, 2, 3, and 4 quarter horizons are highest during this decade. This corresponds nicely with the narrative of Ramey (2009) given in Figure 5. 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 the information parameter, $\alpha_1^G$, 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 of Section 2 to the moving-average coefficients in the tax (capital and labor) and government spending rules [see (26), (27), and (28)].

As an illustrative example of this mapping, consider the following moving-average representation for tax rates

$$\tau_t = \varepsilon_{t-1} - \theta \varepsilon_t. \quad (34)$$

If $|\theta| < 1$, then (34) 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, $\{\varepsilon_{t-j}\}_{j=0}^\infty$.\(^{18}\)

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 (34), which is given by flipping the root, $\theta$, outside of the unit circle

$$\tau_t = \bar{\varepsilon}_t - \theta \bar{\varepsilon}_{t-1} \quad (35)$$

$$\bar{\varepsilon}_t = \left[ L - \theta \right] \varepsilon_t. \quad (36)$$

\(^{18}\)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.
Representation (35) is the Wold representation where $\tilde{\varepsilon}_t$ is the one-step-ahead forecast error associated with forecasting $\tau_t$ conditional on $\{\tau_{t-j}\}_{j=1}^{\infty}$. This representation shows that current and past $\tau_t$ span an equivalent space to current and past $\tilde{\varepsilon}_t$, which is a strictly smaller space than $\varepsilon_t$. To prove this, note that using the Wiener-Kolmogorov optimal prediction formula yields the variance of the one-step-ahead forecast error using representation (35),

$$
E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2 = E\{(L - \theta)\varepsilon_{t+1} - L^{-1}[(1 - \theta L) - 1]\tilde{\varepsilon}_t\}^2
$$

$$
= E\{(L - \theta)\varepsilon_{t+1} - (L - \theta)\varepsilon_{t+1} + \varepsilon_{t+1}\}^2
$$

$$
= \text{var}(\tilde{\varepsilon}_{t+1}) = \sigma_{\tilde{\varepsilon}}^2,
$$

(37)

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

Suppose now that agents are able to condition on current and past structural innovations directly. These agents are able to use (34) to forecast next period’s tax rate. The variance of the forecast error for this process is given by

$$
E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^{\infty}]\}^2 = E\{(L - \theta)\varepsilon_{t+1} - L^{-1}[(L - \theta) + \theta]\varepsilon_t\}^2
$$

$$
= \theta^2 E\{\varepsilon_{t+1}\}^2 = \theta^2 \sigma_{\varepsilon}^2.
$$

(38)

Comparing this forecast error variance with (37) shows that the moving-average coefficient, $\theta$, 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. Therefore, calibrating $\theta$ is tantamount to calibrating agents’ foresight intensity.

Recall from Section 2 that the contemporaneous risk-adjusted implicit tax rate, $\tau_t^{RI}$, is the weighted sum of future tax rates. Therefore, we are able 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

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

(39)

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, $\alpha_1^2$, and the MA coefficient, $\theta$, is determined by the following equality

$$
\frac{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}, \{\tau^{RI}_{t-j}\}_{j=0}^{\infty}]\}^2}{E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^{\infty}]\}^2} = \frac{E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^{\infty}]\}^2}{E\{\tau_{t+1} - E[\tau_{t+1}|\{\tau_{t-j}\}_{j=0}^{\infty}]\}^2}.
$$

$$
1 - \alpha_1^2 = \theta^2
$$

(40)

Equation (40) 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, \( \alpha_1^T \rightarrow 1 \) and \( \theta \rightarrow 0 \), implying perfect one-step-ahead foresight (i.e., (38) goes to zero). If there is no additional information that helps in predicting future taxes beyond the contemporaneous tax rate \( (\alpha_1^T = 0) \), then \( \theta = 1 \) and (34) becomes a fundamental moving average representation. Under this parameter setting, (38) 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 \( \varepsilon_t \) and the current tax rate \( \tau_t \), 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 \( \alpha_1^T \) with \( \alpha_1^G \)). Therefore, estimates of the information parameters \( \alpha_1^T \) and \( \alpha_1^G \) will pin down the foresight intensity—the reduction in the forecast error variance due to fiscal foresight.

The simple example above assumes only one-period of foresight and a constant information parameter. These are unrealistic given the empirical evidence in Section 2 from municipal bonds with one- and five-year maturities and SPF forecasts of government spending up to five quarters ahead, and the evidence that \( \alpha_1 \) varies over time. However, we can proceed with the calculations in (40) to pin down the MA coefficients, \( \theta^K_i, \theta^I_i \) and \( \theta^G_i \) for \( i = 1, \ldots, q \), where \( q \) is the foresight horizon. For example, to calibrate the high degree of tax foresight \( (q = 8) \), we first linearly interpolate the semi-annual estimates of \( \alpha_1^T_t \) to get quarterly data.\(^{19}\) We then take 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 \( \alpha^T_{1,t}^{1980s} \) for \( t = 1980, \ldots, 1990 \). 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 \( \alpha_1^T \)'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, \( \alpha^T_{1,T-j}^{1980s} \) where \( T \) is the date of implementation of tax event and \( j \) is the period of foresight, \( j = 1, \ldots, q \).

As a specific example, the Tax Reform Act of 1986 took effect January 1, 1987. The average of the information parameter \( \alpha_1^T \) estimated from the one- and five-year municipal bond in the fourth quarter of 1986 was approximately 0.19, which corresponds to one-quarter of foresight. The information parameter in the third quarter of 1986 was approximately 0.185, which corresponds to two-quarters of foresight, and so on. We calculate these information parameters for each of the tax events of the 1980s and average them to arrive at a sequence of information parameters \( (j = 1, \ldots, 8) \) for the “high news” regime. We do the same for the medium news regime \( (j = 1, \ldots, 4) \) and the low news regime \( (j = 1, 2) \).

The process for obtaining the sequence of information parameters is slightly different for government spending. This is because we have on hand 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 \( \alpha_1^G \) (see Figure 4) from 2000 through 2009 for the one through four-step-ahead forecasts \( (\alpha_{1,j}^{G,2000s} \text{ for } j = 1, \ldots, 4) \). For the medium news regime, we take the time average

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\(^{19}\)Given the stable nature of the estimates across the regimes, 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.
of $\alpha_1^G$ 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, $\alpha_1$, for each foresight horizon $q$. The mapping from the information parameters, the $\alpha_1$’s, to the moving-average coefficients, the $\theta$’s, follows the one-period example derived above, except the algebra becomes more tedious. For example, 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. As with the previous example, the forecast error associated with the information set $\{\tau_{t-j}\}_{j=0}^\infty$ is derived by flipping the zeros outside the unit circle via Blaschke factors so that $\tau_t = (1 - \xi_1 L)(1 - \xi_2 L)\tilde{\varepsilon}_t$, where $\tilde{\varepsilon}_t = (L - \xi_1)(L - \xi_2)/(1 - \xi_1 L)(1 - \xi_2 L)$.

Assuming the representation is non-fundamental, $|\xi_i| < 1$ for $i = 1, 2$, the ratio of the variance of the one– and two-step-ahead forecast errors conditional on observing the taxes vis-à-vis conditioning on the structural shocks are given by

$$E\{\tau_{t+1} - E[\tau_{t+1}|\{\varepsilon_{t-j}\}_{j=0}^\infty]\}^2 = \theta_0^2, \quad E\{\tau_{t+2} - E[\tau_{t+2}|\{\varepsilon_{t-j}\}_{j=0}^\infty]\}^2 = (1 - \theta_1)^2 \quad (41)$$

The moving-average coefficients are uniquely determined by matching the forecast error variance derived above to the corresponding $\alpha_1$ of the same forecast horizon. If, for example, this were the mapping of the low news tax regime, we would set $\theta_0^2 = 1 - \alpha_{1,T-1}^{\tau,1990s}$ and $(1 - \theta_1)^2 = 1 - \alpha_{1,T-2}^{\tau,1990s}$.

Calculating the variance of the forecast errors in (41) for forecast horizons greater than two yields analytically intractable expressions. However, so long as there exists a unique information parameter for each forecast horizon (that is, for $j = 1, ..., q$), then the mapping from the reduced-form estimates of Section 2 into the DSGE model will be unique.

Proceeding in the fashion described in this section yields the following moving-average representations for the tax and government spending processes:

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

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

**Government Spending**

\textsuperscript{20}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)).
We now turn to analyzing how these different processes alter equilibrium dynamics in the model described in Section 3.

5 Implications of Time-Varying Fiscal Foresight

Direct evidence of the degree of fiscal foresight—whether it is from the narrative studies of Romer and Romer (2010) and Ramey (2009, 2011), the regressions of Fortune (1996), or the tax chronology of Yang (2008)—makes it clear that the degree of foresight varies substantially across time. While some fiscal events are almost wholly surprises, others are years in the making [Steigerwald and Stuart (1997), Romer and Romer (2007)]. At odds with this evidence, formal empirical work on fiscal foresight tends to impose time-invariant degrees of foresight, fixing a priori the horizon over which fiscal events are known [Blanchard and Perotti (2002), Mountford and Uhlig (2009), Mertens and Ravn (2008, 2011), Schmitt-Grohé and Uribe (2008)]. This section explores, in the context of the DSGE model, how different assumptions about the degree of fiscal foresight and the nature of the fiscal news processes alter the model’s predictions of the impacts of anticipated fiscal changes. With the exception of the fiscal news processes, parameters are set at the values estimated in Traum and Yang (2010) (see Appendix B).

The goal of this section is to demonstrate how the different information regimes may lead to under-estimating the effects of news in a standard DSGE model. Several papers, dating back at least to Hall (1971) and more recently Mertens and Ravn (2008, 2011), Ramey (2011) and Schmitt-Grohé and Uribe (2008), have acknowledged and shown that foresight matters in both estimated and calibrated models. But the typical assumption is that news is a time-invariant process. This would be an innocuous assumption if the news regimes were sufficiently close. However, if the news regimes are far apart, estimating a model with time-invariant news runs the risk of under-reporting the impact of fiscal foresight due to the time-averaging of news periods. The true effects of fiscal foresight may be masked by averaging high news periods with periods of no or low news. Given that we have on hand a calibration of 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 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. It is well known that models with these frictions outperform models without these frictions.

The frictions serve to smooth out the response 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 7 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 due to the lack of significant frictions in the labor market, the response of labor to the different news regimes is immediate. Qualitative and quantitative differences are large for the first year, but quickly dissipate so that after 10 quarters the differences are negligible. Similarly, turning off these frictions, as shown by Figure 7, leads to very different responses of the news regimes at all horizons. In fact, removing the frictions can lead to an investment boom on impact [Mertens and Ravn (2011)].

Parameter estimates from Schmitt-Grohé and Uribe (2008) show that real frictions in an estimated model with news shocks are even larger than models without news shocks.\footnote{As an example, the posterior mean of the habit formation parameter in Schmitt-Grohé and Uribe (2008) is 0.85, which is much higher than the typical estimate.} The upshot here is that models estimated with time-invariant news may underestimate the effect of news on some aggregate variables at medium horizons.
Figure 7: 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 8: 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).
5.2 Percentage of Non-Savers A parameter that will obviously have an influence on the extent to which time-variation in news (or news shocks in general) will matter is the proportion non-savers in the economy. Quite intuitively, 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. The effects of foresight rely heavily on agents’ ability to intertemporally substitute. Knowledge of a significant increase in labor taxes in the future has little effect for households that operate hand-to-mouth. Figure 8 shows that as a significant fraction of non-savers are added to the economy, the overall response of employment is mitigated due to the inability to intertemporally substitute. 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 8 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 tantamount to having a high fraction of non-savers in the model.

5.3 Government Spending Figure 9 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: investment and consumption fall due to the government absorbing a larger share of goods, while output increases. However with a high degree of foresight, output could fall in period $t$ as agents anticipate a much higher increase in government consumption in periods $t + 3$ and $t + 4$ (high foresight). A similar result holds with respect to the response of labor; anticipated large increases in government spending in the near future causes 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. Given the now well-known problems with estimating multipliers using fiscal VARs in the presence of foresight [Ramey (2011), Leeper et al. (2011)], DSGE models that model the information process generating foresight explicitly yield more accurate estimates of the true multiplier. However, this is only true if the DSGE model accurately models the news process. Ignoring the 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 9 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 exists 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:

\[
\text{RGFED}_{\text{BY}}^Q = \left( \frac{\text{QI}_{\text{CY}}^Q}{\text{QI}_{\text{BY}}^A} \right) \text{NGFED}_{\text{BY}}^Q,
\]

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. 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

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22 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.
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

\[
NIPD_{CY}^{Q} = \frac{OIPD_{CY}^{A}}{OIPD_{NBY}^{A}},
\]

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 5. 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 6: New Keynesian Model Parameters

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