The Information Content of News Announcements*

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Abstract

This paper investigates high frequency movements of the yield curve around macroeconomic announcements by combining event studies and a no-arbitrage affine term structure model in a new Keynesian model with partial (or imperfect) information. I show that the model fits bond yields and macroeconomic announcement surprises well. The model can fit the empirical responses of bond yields to surprises and the standard deviations of the bond yields around the announcement days. The decomposition of long term nominal zero coupon bond yields shows that the high frequency variation in the long term bond yields is due to a combination of changes in the term premium and revisions to expected short rates. In particular, changes in term premia are as volatile as revisions to expected future short rates. The model estimates imply that around macroeconomic announcement days, average expected short rates and term premia are correlated around announcements. I show that the model implied term premium estimates are strongly correlated with estimates of different reduced form models. The model implies that the behavior of the long term rates in the conundrum period can be explained by a lower term premium. I show that most announcement responses of bond yields are due to changes in the expectations about the output gap.

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1 Introduction

It is well known that bond yields sharply change around macro announcements. For instance, on the days of the employment release, interest rate volatility is about 40-60% higher than on other days (Bauer (2015)). Academics, policymakers and financial market participants pay close attention to bond yield responses at the time of macro announcements for two reasons. First, scheduled releases of macroeconomic indicators constitute an important source of public information about key macro variables to the participants in financial markets. Second, bond yield responses around macro announcements are valuable to understand the link between macroeconomic fundamentals and bond yields. Furthermore, these releases are the only information arriving to the market at the time of the announcement.

In this paper, I examine jumps in bond yields around macro announcements using a structural macro model with time varying term premia. I ask whether these jumps can be reconciled with a simple structural macro model and if so can the model shed some light on why the bond yields move around the macro announcements. Using the macro model, I examine how much of the variation in the long term bond yields around announcements can be attributed to revisions to average expected short rates or changes in term premia. To the best of my knowledge, this is the first paper that analyzes the effects of macro announcements on bond yields and their components using a structural macro model.

Reconciling macro models and the bond yield responses around announcements is an important direction. On the announcement days, financial market participants, policymakers and financial press interpret the changes in bond yields using some underlying model. To make this claim more concrete, I give an example from the financial press. On November 7, 2014, change in nonfarm payrolls was lower than expected. in response, bond yields fell and the Wall Street Journal interpreted this response as a fall in expected future short rates without commenting on the changes in the term premia.[1] This sort of interpretation is not unique to the financial press; similar examples can be found in policy circles and financial institutions. In this paper I check the consistency of these interpretations with a standard macro model and offer a tool to central banks and Wall Street to examine bond yield responses around announcements.

I base the analysis on four building blocks. First I posit a simple new Keynesian model to describe the macro dynamics. New Keynesian models are a prominent tool for policymaking in central banks (Edge and Gürkaynak (2010)). Even though the model in this paper is much simpler than the ones used in policymaking, it will capture the key elements necessary to examine the relationship between bond yields and the macroeconomy. If investors believe that the central bank makes policy using a new Keynesian model (or a model with similar features), it is reasonable to think that they use a similar model to interpret bond yield responses.

Second, one has to modify a standard new Keynesian model in order to incorporate macro announcements. The classic assumption in the macro literature is that agents know the true state of the economy. If agents had perfect information about the state, macro announcements would be informative only if, say, households forget about their income in the previous period since announcements are news about the past. To overcome this problem, I assume that the agents cannot perfectly observe the state of the economy following Svensson and Woodford (2003). Hence the announcements are treated as a source of information about the unobserved state of the economy. With this modification, the model exploits the lumpy manner in which the new information is released to the financial markets at predetermined times through macro news.

Third, I link macro announcements to bond yields. For this purpose, I incorporate term structure into the macro model using a no-arbitrage affine term structure model

[1] The exact quote is: " Treasury bonds strengthened on Friday as a smaller-than-expected increase in October nonfarm payrolls bolstered expectations that the Federal Reserve would be in no hurry to raise official interest rates." - WSJ, 11/7/2014

framework. The idea behind this modification is to jointly model bond yields and macro news and generate the necessary jump in bond yields around announcements times. Then, I decompose long term bond yields into average expected short rates and term premia. This result of this decomposition has to be consistent with a key empirical fact; term premia should be large and variable (Ang and Piazzesi (2003), Kim and Wright (2005), Duffee (2012)). Without this empirical fact satisfied, we cannot have a better understanding of why the bond yields move the way they do around announcements.

Fourth, I estimate the model using macro news and financial market data sampled at high frequency. Aside from the fact that we need high frequency data for the analysis, it has important advantages over low frequency (monthly or quarterly) data. As Gürkaynak and Wright (2013) point out, macro literature is in a state of observational equivalence. That is, macro models with different microfoundations and different policy implications can fit the macroeconomic stylized facts (at low frequency) equally well. Using high frequency data can help researchers to distinguish models with reasonable fit to the high frequency financial and macro data from others that do not fit at all.

I ask two questions about the model. First I ask if the model is consistent with financial market data and, if so, what interpretations we can ascertain from the model regarding the bond yield responses around announcements. To answer these questions, I first estimate the model using Bayesian methods and show that the model fits to real GDP growth, core CPI inflation and nonfarm payroll surprises and bond yields at various maturities quite well. I compare actual and model implied responses of the bond yields to macroeconomic surprises and show that the model can match the average response of yield curve and the volatility of yield changes on announcement days.

Using the estimated model parameters, I decompose the five year, seven year and ten year nominal zero coupon bond yields and show that the decline in long term rates between 1990 and 2007 can be attributed to a combination of lower expected short rates and lower term premia. I show that model implied term premia are highly correlated with the estimates coming from reduced form models. The variance decomposition of the long yields shows that around macroeconomic announcements, average expected short rates and term premia are equally volatile and correlated. Specifically, around real GDP growth and non-farm payroll announcements they are positively correlated. On the other hand, around core CPI announcements, the correlation is positive in short maturities but negative for long maturities. The empirical analysis shows that average expected short rates respond more then term premia to real GDP growth and non-farm payroll announcements. However, term premia respond more than expected short rates to core CPI announcements. Using the estimated model, I analyze the relative contribution of the changes in macroeconomic fundamentals (such as output gap, inflation and inflation target) to responses of bond yields to announcements. I show that bond yields respond mostly to changes in output gap around announcements. Furthermore, I show that changes in macroeconomic fundamentals around announcements cause substantial adjustments in term premia and average expected short rates. However, these magnitude and direction of these adjustments crucially depend on what type of fundamental is being changed around a specific announcement. This result implies that bond yield responses cannot be interpreted as just revisions to expected future short rates, but changes in the term premium has to be part of the story. Finally, I look at the "conundrum" period. Consistent with the literature, the model shows that the fall in the ten year yield during this period can be attributed to a fall in the term premium.

The paper is organized as follows: In Section 2 I present and explain the macro-finance model with the imperfect information that will be used in the paper. In section 3 I describe the data and the estimation methodology used to estimate the model. In section 4 I show the results. Then I conclude in section 5.

2 A New Keynesian Model with Imperfect Information

In this section, I present a standard three-equation new Keynesian model with imperfect information (*à la* Svensson and Woodford (2003)) featuring IS, Phillips curve and short rate equations.

The households maximize the discounted sum of expected utility with the following one-period utility function:

$$U(C_t, N_t, H_t) = \frac{(C_t H_t)^{(1-\gamma)}}{1-\gamma} Q_t - \frac{N_t^{(1+\varphi)}}{1+\varphi}$$

where C_t is consumption, N_t is labor supply, Q_t is the preference shock and H_t is the habit formation. Here γ is the relative risk aversion, φ is the elasticity of labor supply.

An explanation of why habit formation and preference shocks are needed is warranted. As the new Keynesian literature shows, canonical macro models fail to match the well documented persistence of output. This shortcoming is remedied by including habits to the utility function. Fuhrer (2000) shows that a model without habit formation is strongly rejected because the model without habits cannot generate the hump shape in the empirical impulse responses of real output to various shocks. Following Fuhrer (2000), I assume that $H_t = C_{t-1}^{\eta}$, where η is the degree of habit formation.

However, adding endogenous persistence to output may not be enough to match the time series properties of the macro variables. One way to solve this problem is to incorporate demand shocks. One particular example is adding preference shocks to models as a source of variation in the macro observables. It is shown in various studies that these shocks can explain a large fraction of the variation in output and inflation (see Baxter and King (1991), Ireland (2004), Basu and Bundick (2012)).

Following Gallmeyer et al. (2005), I define the growth rate of the preference shock as:

$$-\Delta q_{t+1} = 0.5\,\mathrm{var}(\Delta q_{t+1}) + (\phi_1 b_t)(y_{t+1} - E_t y_{t+1}) + (\phi_2 b_t)(\pi_{t+1} - E_t \pi_{t+1})$$

where $\Delta q_{t+1} \equiv \log(\frac{Q_{t+1}}{Q_t})$ and b_t follows an exogenous AR(1) process. The first term on the right hand side makes sure that $E_t[\frac{Q_{t+1}}{Q_t}] = 1$, that is Q_{t+1} follows a martingale process.

The process for the preference shock defined above is nonstandard in the macro literature. Studies that incorporate preference shocks as a source of variation model them as highly persistent autoregressive processes. Different from the literature, I assume that the preference shocks follow a random walk and they depend linearly on output and cost push shocks. As I show below, this nonstandard specification helps to match the financial data without compromising the macro fit of the model where the standard preference shock specifications are only useful in matching the macro variables.

The implication of the preference shock for the term structure is that it helps the model to generate time varying term premia. It is well known in the macro-finance literature, that a macro model with nontrivial term premia requires either time varying price of risk or time varying quantity of risk. In this paper, I introduce time varying price of risk using the preference shock specification. Hence the model generates time varying term premia without resorting to time varying quantity of risk, which requires time variation in the second moments incorporated through a third order approximation (Rudebusch and Swanson (2013)). The details of why this specification implies time varying risk premia is given in the next subsection.

Given the instantaneous utility function, households choose how much to consume and how much labor to supply with respect to the following budget constraint:

$$\int_0^1 P_t(j)C_t(j)dj + e^{-i_t}B_t \leq B_{t-1} + W_t N_t + T_t$$

where B_t is the amount of one period bonds purchased by the household with a price of e^{-i_t}, $P_t(j)$ is the price of good j, T_t is the lump sum transfer and W_t is the wage earned by the household.

There is a continuum of monopolistically competitive firms producing differentiated goods indexed by j. The firms have identical Cobb-Douglas production functions linear in labor:

$$Y_t(j) = A_t N_t(j)$$

where A_t is the aggregate technology which follows an AR(1):

$$\log(A_t) = \rho \log(A_{t-1}) + \varepsilon_t^a$$

$$\varepsilon_t^a \sim N(0, \sigma_a^2)$$

Firms set prices according to Calvo mechanism where a fraction θ of firms reset prices in a given period. Of the firms that reset prices, a fraction $(1-\omega)$ optimize their price decision and take into account that their price may be effective for more than one period while a fraction ω of price setters use a "rule of thumb" as in Gali and Gertler (1999). The "rule of thumb" firms set their price equal to last period's average reset price plus the lagged inflation rate. As Gali and Gertler (1999) shows, backward indexation in prices is an important component of new Keynesian models to match the inertial response of inflation to economic shocks.

The log-linearized model is given by the IS and Phillips curves, respectively:

$$y_t = \frac{\gamma}{\gamma - \eta + \gamma\eta} E_t y_{t+1} - \frac{\eta(1-\gamma)}{\gamma - \eta + \gamma\eta} y_{t-1} - \frac{1}{\gamma - \eta + \gamma\eta} [i_t - E_t \pi_{t+1} + E_t \Delta q_{t+1}] + \varepsilon_t^y$$

$$\pi_t = \frac{\beta\theta}{\theta + \omega(1 - \theta(1-\beta))} E_t \pi_{t+1} + \frac{\omega}{\theta + \omega(1 - \theta(1-\beta))} \pi_{t-1} + \frac{(1-\omega)(1-\theta)(1-\theta\beta)}{\theta + \omega(1 - \theta(1-\beta))} mc_t + \varepsilon_t^\pi$$

where y_t, π_t and i_t are real output, inflation and short rate at time t and $\Delta q_{t+1} = \log(Q_{t+1}) - \log(Q_t)$. The real marginal cost at time t, given by mc_t, can be found by dividing the real wage by the productivity. The log-linearized marginal cost is given by:

$$mc_t = (\gamma + \varphi)y_t + (\eta - \eta\gamma)y_{t-1} - q_t - (1 + \varphi)a_t$$

where the relationship $n_t = y_t - a_t$ is used to substitute out labor supply. The IS and Phillips curves in this model are very standard in the new Keynesian literature with preference shocks (see Ireland (2004) for an example).

Given the set up of the economy above, behavior of monetary policy will close the model. I assume that the monetary policy controls the short rate by minimizing the following quadratic loss function:

$$L_t = E_t \left[\sum_{k=0}^{\infty} \beta^k [\lambda_y (y_{t+k} - \bar{y}_{t+k})^2 + (\pi_{t+k} - \pi^*_{t+k})^2 + \lambda_i (i_{t+k} - i_{t+k-1})^2 \right]$$

This loss function implies that the central bank tries to minimize the deviations of output level from its potential ($\bar{y}_t$), deviations of inflation from its time-varying long run level (π^*_t) and does interest rate smoothing. The relevant λs are the respective weights of these objectives. Potential output is given by:

$$\bar{y}_t = \frac{\eta(1-\gamma)}{(\varphi+\gamma)} y_{t-1} + \frac{(1+\varphi)}{\varphi+\gamma} a_t + \frac{1}{(\varphi+\gamma)} q_t$$

Following Del Negro and Eusepi (2011), I assume that the time-varying long run inflation follows an exogenous stationary AR(1) process:

$$\pi^*_t = \rho_{\pi^*} \pi^*_{t-1} + \varepsilon^{\pi^*}_t$$

$$\varepsilon^{\pi^*}_t \sim N(0, \sigma^2_{\pi^*})$$

Although the process for long run inflation chosen to be stationary, my prior for ρ_{π^*} will push this process to be highly persistent. As I show below, this process will be helpful in matching both the persistence and the decline in long term inflation expectations in my estimation sample.

In this paper, I am not making any claims about the optimality of this loss function. One can derive the functional forms for λs that are represented by the structural parameters of the model. Instead, this loss function should be interpreted as an approximation to the behavior of the central bank (Svensson (2008)).

In a model with perfect information, where the agents observe the true state of the economy, the model solution can always be expressed as:

$$X_t = \Gamma X_{t-1} + B\varepsilon_t$$

where X_t is the vector of predetermined (or state) variables and ε_t is the vector of structural shocks. In this model, the decision rules of the agents are functions of the true state of the economy since the true state of the economy is fully observable.

However, I am relaxing this assumption and assume that the agents cannot observe the true state of the economy. The central bank and the households observe a vector of observable variables at time t. Using the new information embedded in this vector of observables, agents estimate the state of the economy. Contrary to the perfect information case, the solution to this filtering problem makes decision rules depend on the agents' best estimate of the true state of the economy, which can be represented as:

$$\begin{bmatrix} X_t \\ X_{t|t} \end{bmatrix} = \Phi \begin{bmatrix} X_{t-1} \\ X_{t-1|t-1} \end{bmatrix} + \Psi \begin{bmatrix} \varepsilon_t \\ v_t \end{bmatrix}$$

where $X_{t|t}$ is the optimal estimate of the state and v_t is the vector of one-step ahead prediction errors. The reason why one-step ahead prediction errors show up in the solution is as follows. As of time $t-1$, agents predict the state vector using the information set that is available to them as of $t-1$. When the new information is available to the agents as of time t, the prevailing state might be different from what the agents have predicted. To take that into account, the solution should incorporate one step ahead prediction errors as a source of disturbance to the model.

The final stage of fully explaining the solution of the model is to describe how $X_{t|t}$ evolves. The procedure of estimating $X_{t|t}$ follows the standard Kalman updating scheme. This procedure requires a measurement equation and a transition equation. The measurement equation will determine how the observable variables are related to the state vector. The transition equation will be the solution of the model under imperfect information.

At a given time t, agents observe an announcement or a survey along with a vector of bond yields associated with the observed announcement or survey. These two different types of observables are related to the state of the economy through different measurement equations. The measurement equation for the announcements and surveys is given by:

$$Z_t^{j,m} = DX_t + u_t^{j,m}$$

where $Z_t^{j,m}$ is the macroeconomic variable observed at time t related to topic j of type m, D is the selection vector consisting of zeros and ones to pick the relevant variable from the true state vector and $u_t^{j,m}$ is the noise. The left hand side of this equation is generally a scalar; most of the time agents can observe a single announcement or a survey. In this measurement equation, I allow the variances of the noise to vary with topic j and type m. Specifically, I allow the surveys to have a greater noise than the announcements. Hence the survey expectations will have less information content about the state of the economy relative to announcements due to the additional noise in the measurement equation.

The measurement equation for the bond yields is given by:

$$\mathcal{Y}_t^O = \mathcal{Y}_t + u_t^{\mathcal{Y}}$$

where $\mathcal{Y}_t^O$ is the vector of bond yields observed at time t, $\mathcal{Y}_t$ is the vector of true bond yields and $u_t^{\mathcal{Y}}$ is the noise vector. In this model, true bond yields (linearly) depend on the true and estimated state of the economy. This implies that the yield curve is not fully revealing, that is the agents cannot back out the true state of the economy by looking at the yield curve. Under this assumption, both the econometrician and the agents in the model learn about the state of the economy.

Using the measurement and transition equations (and assuming that all agents have the same information set), we can describe the process for estimating the state of the economy. To fix ideas, suppose that news arrives at time t. Agents compare their one-step ahead predictions of the observables (generated from one-step ahead predictions of the state) with their realizations. There will be a discrepancy between the realizations and the predictions made for announcements/surveys and bond yields in period $t-1$. Using these prediction errors, agents linearly update their beliefs about the state of the economy:

$$X_{t|t} = X_{t|t-1} + K \begin{bmatrix} u_t^{j,m} \\ u_t^{\mathcal{Y}} \end{bmatrix}$$

where K is the Kalman gain. The mathematical details of the updating procedure can be found in the Appendix.

Finally, I will specify how the short rate is related to macroeconomic fundamentals. In this setup, Svensson and Woodford (2003) show that with linear constraints and a quadratic loss function, the optimal short rate can be written as a linear function of the estimate of the state variables:

$$i_t = F X_{t|t}$$

As shown in Svensson and Woodford (2003), in this model the "certainty equivalence" holds. Estimation of the state of the economy and the minimization of the intertemporal loss function can be separated. This separation principle implies that the parameters of the vector F will be equal to the case where the central bank can observe the state perfectly. For future reference, I also define the vector $\bar{F} = \begin{bmatrix} \mathbf{0}_{1\times S} & F \end{bmatrix}$ where S is the size of the state vector. This is simply an augmented F vector with loadings of zeros on the true state vector.

2.1 Implications for the Term Structure

In this section, I explain the implications of the macro model for the term structure.

Under no-arbitrage, the price of a default-free n-period nominal zero-coupon bond that pays \$1 at maturity satisfies:

$$1 = E_t[M_{t+1}R_{t+1}] \tag{1}$$

where $R_{t+1} = \frac{p_{t+1}^{(n-1)}}{p_t^n}$ is the one period return for an n period bond between t and $t+1$ and M_{t+1} is the stochastic discount factor (SDF).

Following the macro-finance literature, I assume that the SDF is implied by the consumption Euler equation. Specifically, the SDF that the agents use to price bonds is the intertemporal marginal rate of substitution given as:

$$m_{t+1} = \log\beta - \gamma\Delta c_{t+1} + (\gamma\eta - \eta)\Delta c_t - \pi_{t+1} + \Delta q_{t+1}$$

The big role of the unobserved preference shock is that, its shocks raise the volatility of marginal utility. This is an important feature of the model because the asset pricing literature finds that the volatility of marginal utility in power utility models is much too low to explain the average risk premia on assets (Duffee (2012)). In the given specification, preference shocks are linear in output and inflation shocks which are assumed to be the intuitive sources of risk that the investors fear. Moreover, the sensitivity of $-\Delta q_{t+1}$ to output and inflation shocks is time varying due to the latent factor b_t. A shock that would increase b_t will make m_{t+1} more responsive to a given structural shock where as a lower b_t will make it smoother. This difference in volatility of marginal utility will make the conditional covariance between m_{t+1} and the price of the bond time varying, implying a time varying risk premium.

Using the no arbitrage condition and assuming that the SDF is conditionally log normally distributed, we can rewrite m_{t+1} in a general form:

$$m_{t+1} = (-i_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t\varepsilon_{t+1})$$

where λ_t is the time varying price of risk which is affine in the state of the economy $\bar{X}$:

$$\lambda_t = \lambda_0 + \lambda_1\bar{X}_t$$

For the derivation of the time varying price of risk, I use the shocks to m_{t+1} represented as:

$$m_{t+1} - E_t m_{t+1} = -\lambda_t'\varepsilon_{t+1}$$

Given that the price of risk λ_t is an affine function of the state:

$$m_{t+1} - E_t m_{t+1} = -(\lambda_0 + \lambda_1\bar{X}_t)'\varepsilon_{t+1}$$

where ε_{t+1} is the vector of structural shocks.

Using the model implied equation for m_{t+1} I can rewrite the left hand side as:

$$m_{t+1} - E_t m_{t+1} = \gamma(y_{t+1} - E_t y_{t+1}) + (\pi_{t+1} - E_t\pi_{t+1}) - (\Delta q_{t+1} - E_t\Delta q_{t+1})$$

Rearranging and using the assumed functional form for the preference shock I have:

$$m_{t+1} - E_t m_{t+1} = (\gamma - \phi_1 b_t)(y_{t+1} - E_t y_{t+1}) + (1 - \phi_2 b_t)(\pi_{t+1} - E_t\pi_{t+1})$$

The model solution implies that I can write forward looking variables in the model (output and inflation) as a function of the state variables:

$$\begin{bmatrix} y_{t+1} \\ \pi_{t+1} \end{bmatrix} = \bar{G}\bar{X}_{t+1}$$

where $\bar{X}_{t+1} = \begin{bmatrix} X_{t+1} & X_{t+1|t+1} \end{bmatrix}'$ is the state vector and $\bar{G} = \begin{bmatrix} G & G - G^1 \end{bmatrix}$. Using the above equality and the VAR(1) form for the state vector, the shocks to m_{t+1} can be written as:

$$m_{t+1} - E_t m_{t+1} = \begin{bmatrix} \gamma & 1 \end{bmatrix} \bar{G}_y \Sigma \varepsilon_{t+1} + b_t \begin{bmatrix} -\phi_1 & -\phi_2 \end{bmatrix} \bar{G}_\pi \Sigma \varepsilon_{t+1}$$

$$m_{t+1} - E_t m_{t+1} = -(\begin{bmatrix} -\gamma & -1 \end{bmatrix} \bar{G}_y \Sigma + b_t \begin{bmatrix} \phi_1 & \phi_2 \end{bmatrix} \bar{G}_\pi \Sigma) \varepsilon_{t+1}$$

where $\bar{G}_y$ and $\bar{G}_\pi$ are the rows of matrix $\bar{G}$ that are associated with output and inflation.

This equality gives λ_0 and λ_1:

$$\lambda'_{0,b} = \begin{bmatrix} -\gamma & -1 \end{bmatrix} \begin{bmatrix} \bar{G}_y \\ \bar{G}_\pi \end{bmatrix} \Sigma$$

$$\lambda'_{1,b} = \begin{bmatrix} \phi_1 & \phi_2 \end{bmatrix} \begin{bmatrix} \bar{G}_y \\ \bar{G}_\pi \end{bmatrix} \Sigma$$

where $\lambda_{0,b}$ and $\lambda_{1,b}$ correspond to the rows of λ_0 and λ_1 that are associated with b_t. Conditional on b_t being positive, if the parameters ϕ_1 and ϕ_2 are positive, a positive shock to output corresponds to an unexpectedly low marginal value of consumption. The bond prices go up when the marginal utility of consumption is high, implying positive risk premium.

I assume that the natural log of the bond prices is affine in state of the economy:

$$\log p_t^{(n)} = A_n + B_n \bar{X}_t$$

where $\bar{X}_t = [X_t, X_{t|t}]'$. The coefficients A_n and B_n follow the following recursions:

$$
\begin{aligned}
A_n =& A_{n-1} - B_{n-1}(\Psi'\Sigma\Sigma'\lambda_0) + 0.5B'_{n-1}\Psi'\Sigma\Sigma'\Psi B_{n-1} - \delta_0 \\
B_n =& B_{n-1}(\Phi - \Psi'\Sigma\Sigma'\lambda_1) - \delta_1
\end{aligned}
$$

where Σ is the variance-covariance matrix of the structural shocks and theoretical prediction errors and

$$
\begin{aligned}
\delta_0 =& -\log\beta + 0.5\lambda_0\Sigma\lambda_0 \\
\delta_1 =& \bar{F}
\end{aligned}
$$

Then the continuously compounded nominal yield to maturity on the n-period nominal zero-coupon bond is given by:

$$
y_t^{(n)} = -\frac{1}{n}(A_n + B_n\bar{X}_t)
$$

3 Data and Estimation

In this setup, the central bank and the investors can observe macro variables with noise and a lag. The macro variables that economic agents observe are real GDP growth, core CPI inflation and nonfarm payroll announcements and surveys, including the five to ten year (median) inflation expectations from the Blue Chip. In the estimation non-farm payrolls proxy the hours worked. Core CPI inflation and non-farm payroll announcements/surveys are sampled monthly, real GDP growth announcements/surveys are sampled quarterly and Blue Chip forecasts are sampled biannually. Given this difference in sampling frequencies, one has to make sure that the lower frequency observables are consistent with the underlying model's frequency, which is monthly. In particular, I have to make sure that the real GDP growth announcements and surveys and Blue Chip forecasts are consistent with the timing of the state variables of the model. The model implied quarterly real GDP growth rate can be calculated as:

$$\Delta y_t^Q = \sum_{i=t-3}^{t-1} y_i - \sum_{i=t-6}^{t-4} y_i \tag{2}$$

where y_t is the model implied level of output at time t and Δy_s^Q is the model implied real GDP growth rate in the previous quarter.

To make sure that the inflation expectations are consistent with the model's frequency, I estimate a state space model (using Kalman filter) that allows me to fill in the missing observations in the Blue Chip forecasts at a monthly frequency. Specifically, I estimate the missing values for the Blue Chip forecasts by extracting a factor from the long run inflation expectations from the Michigan Survey of Households. Figure 1 shows the interpolated and observed Blue Chip forecasts. The interpolated series will be used for estimation.

I estimate the state-space system using Bayesian methods and the Kalman filter. The priors for the structural parameters are given in Table 1. I estimate the model only on the announcement/survey or inflation expectation release days starting from January 1, 1990 until December 31, 2007 because of the zero lower bound. I assume that the survey expectations of the real GDP growth, core CPI and non-farm payrolls are released the day before their respective announcements.

Finally, I assume that the central bank and the investors can observe one month, one year, two year, five year, seven year and ten year nominal zero coupon bond yields at the daily frequency[2].

The estimated model will be used to decompose the long term bond yields into average expected short rates and term premia. To do the decomposition I will use the fact that

[2]One month bond yields are calculated by Le and Singleton using Fama-Bliss approach. I would like to thank Ahn Le and Greg Duffee for sharing this data. Other bond yields are from Gürkaynak, Sack and Wright (2007) available at the daily frequency.

the n-period zero coupon bond yield can be written as:

$$y_t^{(n)} = \frac{1}{n}\sum_{j=0}^{n-1} E_t(i_{t+j}) + tp_t^{(n)}$$

where $\frac{1}{n}\sum_{j=0}^{n-1} E_t(i_{t+j})$ is the average expected short rate in n-periods and $tp_t^{(n)}$ is the term premium of the n-period bond. The average expected short rates are calculated by first forecasting the short rate n-periods into the future at every t and taking the average of these forecasts.

4 Results

I report the posterior estimates of the model parameters in Table 1. The priors and posterior estimates of the measurement error standard deviations of real GDP growth and inflation expectations are shown as annualized percentages where as the non-farm payroll measurement error standard deviations are in thousands. I keep core CPI inflation measurement error standard deviation as monthly to be consistent with the data. For ease of interpretation I report the bond yield measurement errors in annualized basis points.

I choose fairly tight but rather standard priors for the structural parameters of the model. One exception to this are the priors for ϕ_1 and ϕ_2 which affect the time varying price of risk. Since the price of risk is unobservable and the estimates change from model to model, I do not have a good guidance on how to choose the priors. Due to this limitation, I choose flat priors for ϕ_1 and ϕ_2 with positive support and let the data decide on their values. Similarly, I choose flat priors for the standard deviation of the structural shocks except for the inflation target and the latent factor shocks. The reason for this is to make sure that the inflation target and b_t are not too volatile given that the interest rates and inflation expectations are very smooth in the data.

The standard deviations of the theoretical prediction errors for real GDP growth, core

CPI inflation, non-farm payrolls and inflation expectations are denoted by $\sigma_y^{v,cb}$, $\sigma_\pi^{v,cb}$, $\sigma_n^{v,cb}$ and $\sigma_{\pi^*}^{v,cb}$, respectively. To limit the number of parameters to estimate, I assume that the standard deviations of the theoretical and empirical measurement errors for the bond yields are equal. I have flat priors for the theoretical and empirical measurement errors for real GDP growth, core CPI inflation and non-farm payrolls but tight priors with low prior mode for inflation expectations. Similarly I choose flat priors for the survey noise standard deviations.

The posterior estimates for the structural parameters are similar to the estimates in the literature. Posterior means for ϕ_1 and ϕ_2 are positive and similar in magnitude, implying that the investors require roughly the same amount of compensation for output and inflation shocks. One striking results is that the posterior means for the bond yield measurement error standard deviations are not more than 10 basis points on annual basis. One important reason for this is the time varying term premia generated by the model. The posterior standard deviations are lower than the priors' suggesting that the data is very informative on the parameters of the model.

Figure 2, shows the model fit for the macro factors on the announcement/survey days of the respective macro variables. The model has a great fit for the macroeconomic observables, especially to announcements and surveys. Furthermore, the model fits inflation expectations very well. This is an important result, since inflation expectations are not fit well using standard macroeconomic models as pointed out by Del Negro and Eusepi (2011). As the posterior estimates suggest, the fit of the model to the bond yields is comparable to various affine term structure models used in the literature (see Kim and Wright (2005) for an example). The fit is shown in Figure 3.

To better understand the model fit, in Figure 4 I plot the model implied surprises against the actual surprises. The surprises are defined as the survey expectations minus the announcements. The correlations are shown to be very high between the model implied and actual surprises implying that the model performs well in matching the surveys and announcements.

Can the model match the average responses of bond yields around announcement days and their respective volatility? To answer this question, I first estimate the responses of the bond yields to macroeconomic surprises using an event study regression. Then I compare these estimates with the model implied ones. This comparison is shown in 5, where the black crosses show the 95% confidence interval for the event study coefficients. The model can match the hump shape of the responses of bond yields. Generally, model implied responses are within the confidence intervals of the regression coefficients.

Next, I compare the model implied and actual volatility curves around nonfarm payroll announcements. Fleming and Remolona (1999) and Piazzesi (2001) show that on employment report days, the level of volatility is higher and hump-shaped in maturity. That is, the most volatile yields on employment report days are the intermediate maturity yields where as the volatility is lower at the short end of the yield curve. As Kim and Wright (2016) point out, this is an empirical regularity that the standard term structure models cannot capture. Following Kim and Wright (2016), I look at the volatility in yield changes caused by employment report announcements. The result is shown in Figure 6. The actual volatility curve is the square root of the difference between the variance of the change in bond yields around nonfarm payroll announcement days and non-announcement days. Model implied volatility curve is the volatility of model implied yield changes around nonfarm payroll announcement days. The figure shows that the actual volatility in yield changes is higher for the 5 to 10 year bonds. The model can capture this fact. Hence the model is successful in matching the hump shaped volatility curve around nonfarm payroll announcement days.

After showing that the model fits the observables well, the next step is to decompose the long term rates into its components. Here I will focus on five year, seven year and ten year nominal zero coupon bond yields and decompose them into average expected short rates in respective maturities and the term premia. Figure 7 shows this decomposition. The estimated model implies that the decline in the long term bond yields is due to lower expected short rates and lower term premia. This is consistent with widely used affine term structure models in the literature (see Kim and Wright (2005)). Even though their estimates of these components differ from model to model, this result is considered an empirical regularity (see Rudebusch (2007)). This is a result that the model in this paper generates by staying within in the confines of a structural macro model with restrictions on the price of risk derived from the model implied SDF.

One natural question is to ask how the model implied term premium is related to other measures of term premium estimated in the literature. To answer that question, I compare the model implied five year, seven year and ten year nominal term premia estimates with the corresponding estimates of Kim and Wright (2005). As Figure 8 shows, these estimates move together with a correlation around 0.9. It is encouraging to see that the macroeconomic model in this paper can generate similar variation in the term premium as one of the widely used reduced form affine term structure models.

With this result in hand, I look at an interesting episode in my sample. Between June 29, 2004 and February 2, 2005 the FOMC increased the target federal funds rate by 1.5 percentage points. During this tightening period, nominal ten-year bond yields declined around 0.7 percentage points. This response of ten-year yields to an increase in the federal funds rate was called a conundrum by then Federal Reserve chairman Alan Greenspan in his February 2015 testimony to Congress. It was a conundrum because historically, increases in the federal funds rate are accompanied by higher long term yields. Backus and Wright (2007) argue that this decline in the long term yields is the result of lower term premia using an no-arbitrage affine term structure model. Rudebusch, Swanson and Wu (2006) use a macro-finance model to analyze the conundrum period and show that this period is associated with lower term premia and relatively stable expected short rates. However, they point out that their model generates large and negative fitting errors for long term rates over the conundrum period relative to the rest of the sample (around 40 basis points) and argue that this period can be explained by other fundamentals that are not modeled. I follow these papers and look closely at the conundrum period to see what my model implies for the behavior of the long term rates over that part of the sample.

I show the decomposition of the ten year bond yields in Figure 9 along with the short rates. During the conundrum period (denoted by the vertical lines), the model implies that the decrease in the ten year yield is solely due to lower term premium. Furthermore, although the ten-year fitting errors are negative, they are not larger than 10 basis points (not shown here). This is a striking difference between the macro-finance model used in Rudebusch, Swanson and Wu (2006). The main reason for this difference is due to the preference shock incorporated in my model. The preference shock might capture changes in the net supply of Treasuries, changing issuance, and especially changing demand from China and other official purchases of Treasuries.

Using the above decomposition of long term rates, I can ask how much of the variation in the changes of long term bond yields around announcement days is due to changes in expected future short rates and term premia. However, given the possible correlation between the revisions to expected short rates and the changes in term premia, I cannot obtain a clean measure of the variance shares. Hence I report the standard deviations of the changes bond yields, average expected short rates and term premia around announcements. The result of this exercise is given in Tables 2, 3 and 4 for five year, seven year and ten year yields, respectively. A quick look at this table reveals that the non-farm payroll announcements are related to higher variation of the bond yield changes around the announcement days, being the most important announcement for the bond yields. Furthermore, expected short rates and the term premia are equally volatile and correlated around announcement times. Specifically, revisions to averaged expected future short rates are positively correlated with changes in term premia around real GDP growth and non-farm payroll announcements, whereas they are slightly negatively correlated around core CPI announcements. However, this negative correlation is very small. Biggest correlation between these components are observed around non-farm payroll announcement times, which is around 0.5.

All in all, above evidence suggests that bond yields change because of revisions to expected future short rates and term premia. Hence lumping all of the change in bond yields to revisions of expected short rates is only half of the story. To make this point more concrete, I show the model implied responses of bond yields and their respective components to macroeconomic surprises in Figure 10. The figure shows that revisions to expected future short rates affect bond yields at all maturities around announcements. Changes in risk premia affect bond yield responses but the magnitude of this effect depends on the announcement type. Specifically, bond yield changes around real GDP growth and non-farm payroll announcements are mainly due to revisions to expected future short rates. On the other hand, long term bond yield changes around core CPI announcements are mainly due to changes in the term premia. Furthermore, revisions to expected future short rates and changes in the term premia move in opposite directions around core CPI announcements for longer maturities.

The structural model posited in this paper enables me to identify the relative importance of the state variables in terms of contributing to the bond yield responses to macroeconomic announcements. To this end, I show the model implied responses of bond yields to output gap, inflation and inflation target around macroeconomic announcements. The responses are given in Figure 11. The responses of bond yields for all announcement groups are due mostly to output gap changes. This effect is strongest around non-farm payroll announcements. However, both changes in inflation and inflation target affect bond yields in all maturities but with a smaller contribution.

How do components of bond yields respond to changes in state variables? The model implied responses of average expected short rates and term premia to state variables are given in Figures 12 and 13, respectively. Responses of expected short rates to state variables depend on the announcement types. Around real GDP growth announcements, responses to changes in inflation dominates the responses to other state variables. These responses show that a better than expected real GDP growth announcement makes agents to revise their short rate expectations upward due to higher inflation in the long run. Similar pattern is observed for the responses to output gap around non-farm payroll announcements. Around core CPI announcements, expected short rates revisions are mainly due to higher expected inflation. Higher than expected core CPI announcement, increases inflation target and hence expected future short rates. Even though responses to output gap and inflation dominate the response of expected future short rates in short maturities, response of the expected short rates to a change in inflation target increases sharply for the two-year maturities and stays elevated for longer maturities.

On the term premia side, we can identify a uniform pattern. Term premia respond mostly to changes in output gap. Furthermore, responses of term premia to output gap are greater than the responses of expected future short rates around real GDP growth and core CPI announcements, implying that the responses of bond yields to output gap can be mainly explained by changes in term premia. Around core CPI announcements, term premia increase with positive changes in inflation whereas they fall with positive changes in the inflation target. Similarly, around real GDP growth announcements, term premia fall with positive changes in inflation. Overall, these results imply that term premia responds to macroeconomic fundamentals hence to macroeconomic announcements.

5 Conclusion

In this paper, I proposed a first step in reconciling the high frequency macroeconomic surprises and the widely-used lower frequency macroeconomic models. To this end, I estimated a standard small scale new Keynesian model with imperfect information and show that it can match the basic properties of macroeconomic surprises and bond yield responses to these surprises. Furthermore, the estimated model can be used to decompose the effects of the announcements into changes in expected short rates and the term premia. This exercise shows that around announcements both of these components are highly variable and have offsetting effects on the bond yield responses. This finding supports the view that the announcements cannot be solely interpreted by their effects on the expected short rates since the term premia would be an important factor determining the response of the bond yields.

However, there are some limitations of the paper that I leave for future work. One of the most pressing issues is the information arrival. In this paper I assumed that the information arrives in a lumpy manner between evenly spaced time intervals. This implies that the variance-covariance matrix of the shocks is independent of time. However a more realistic implementation would be to have a continuous time model that would allow the announcements happening in different days in a given month.

6 Appendix

Following Svensson and Woodford (2003) I can write the model in the following form:

$$\begin{bmatrix} X_{t+1} \\ x_{t+1|t} \end{bmatrix} = \begin{bmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{bmatrix} \begin{bmatrix} X_t \\ x_t \end{bmatrix} + \begin{bmatrix} B_1 \\ B_2 \end{bmatrix} FX_{t|t} + \begin{bmatrix} C \\ 0 \end{bmatrix} \varepsilon_{t+1}$$

where X_t is the vector of predetermined or state variables at time t, x_t is the vector of forward looking variables at time t and ε_{t+1} is the vector structural shocks. Specifically, the predetermined variables of the model are:

$$X_t = \begin{bmatrix} a_t & y_{t-1} & \pi_{t-1} & n_{t-1} & \varepsilon_t^y & \varepsilon_t^\pi & \pi_t^* & \pi_{t-1}^* & \varepsilon_t^* & i_{t-1} \\ \Delta i_{t-1} & y_{t-2} & y_{t-3} & y_{t-4} & y_{t-5} & y_{t-6} & b_t & q_t & & \end{bmatrix}'$$

The dynamics of the model can be summarized as:

$$\begin{aligned} X_t =& HX_{t-1} + JX_{t-1|t-1} + C\varepsilon_t \\ x_t =& G^1 X_t + (G - G^1)X_{t|t} \\ X_{t|t} =& X_{t|t-1} + K(Z_t - LX_{t|t-1} - MX_{t|t}) \\ Z_t =& LX_t + MX_{t|t} + v_t \end{aligned}$$

where Z_t is the vector of variables that are observable to the central bank and K is the Kalman gain. The coefficient matrices G and G^1 satisfy:

$$\begin{aligned} G =& (A_{22} - GA_{12})^{-1}[-A_{21} + GA_{11} + (GB_1 - B_2)F] \\ G^1 =& -(A_{22})^{-1}A_{21} \end{aligned}$$

Given G and G^1, transition matrices H and J can be defined as:

$$\begin{aligned} H \equiv & A_{11} + A_{12}G^1 \\ J \equiv & B_1 F + A_{12}(G - G^1) \end{aligned}$$

The quadratic loss function of the central bank can be written in a matrix form:

$$L_t = Y_t W Y_t'$$

where $Y_t = \begin{bmatrix} y_t - \bar{y}_t & \pi_t - \pi_t^* & i_t - i_{t-1} \end{bmatrix}$ is the vector of control variables and W is a diagonal matrix with associated λs on the diagonal.

The Kalman gain matrix has the usual definition:

$$\begin{aligned} K =& PL'(LPL' + \Sigma_{vv})^{-1} \\ P =& H[P - PL'(LPL' + \Sigma_{vv})^{-1} LP]H' + \Sigma_{\varepsilon\varepsilon} \end{aligned}$$

where P is the variance-covariance matrix of the prediction errors.

Notice that the Kalman gain matrix K is independent of the policy chosen. The estimation of the state and the optimal policy given the state can be treated as separate problems.

The model solution implies the following transition equation:

$$\bar{X}_t = \Phi \bar{X}_{t-1} + \Psi \begin{bmatrix} \varepsilon_t \\ v_t \end{bmatrix}$$

where

$$\Phi = \begin{bmatrix} H & J \\ KLH & H + L + KLJ + K + L(H + J) \end{bmatrix}, \Psi = \begin{bmatrix} C & 0 \\ KLC & K \end{bmatrix}$$

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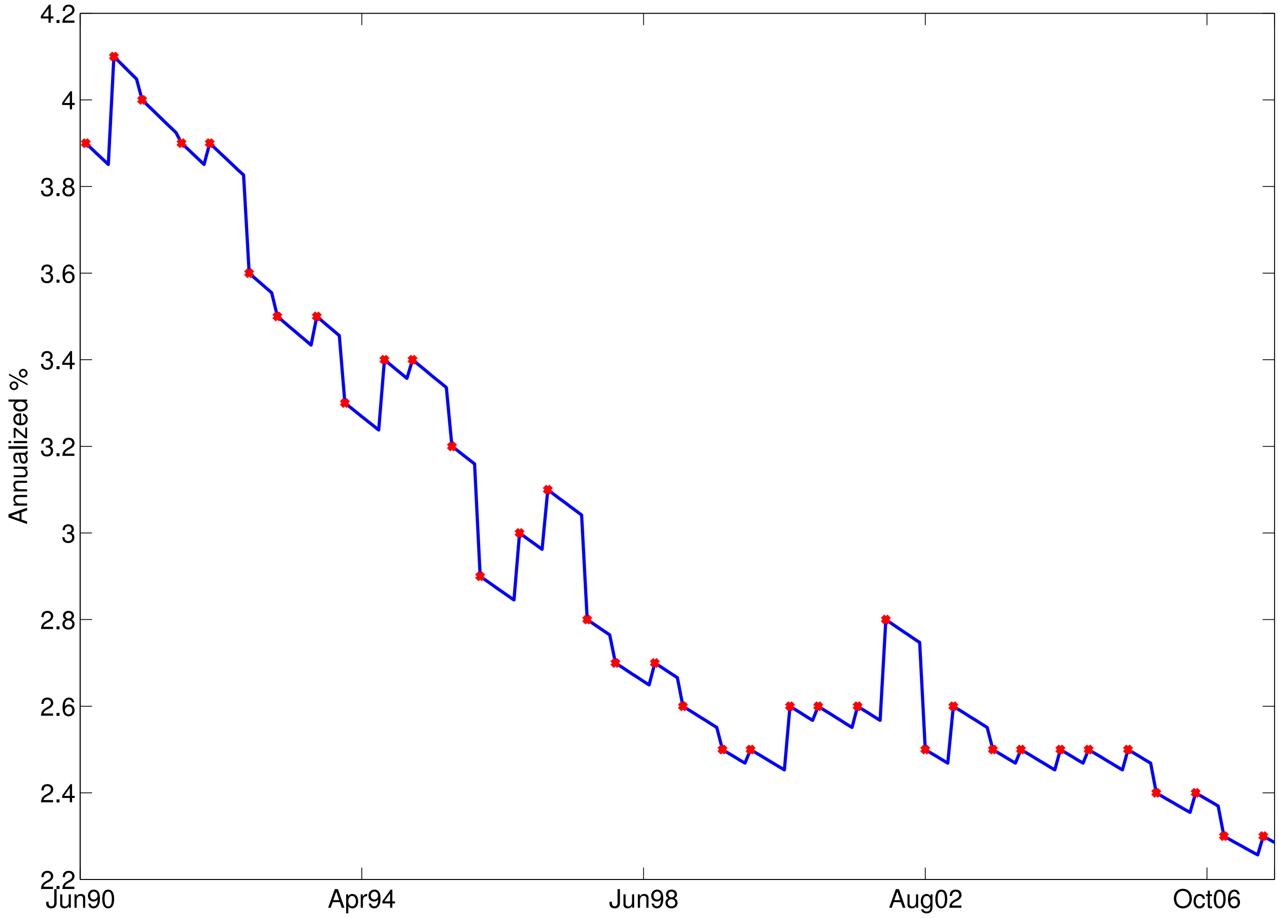

Figure 1: Interpolated vs Actual Blue Chip Long Run Inflation Forecasts

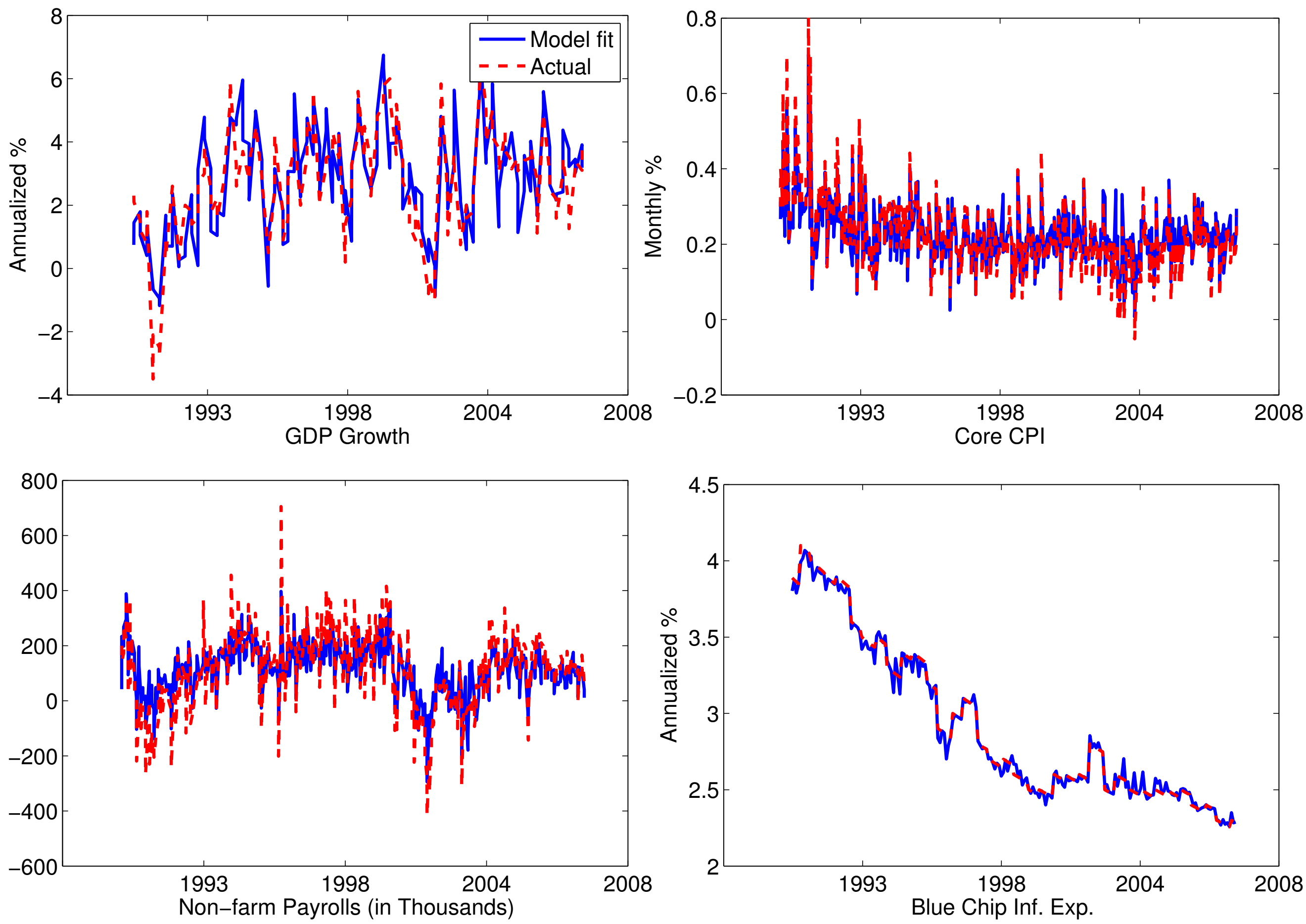

Figure 2: Model fit vs. Data for the Macro Factors. Data for the macro factors are described in the text. The figure shows the announcement/survey days of the respective macro variables.

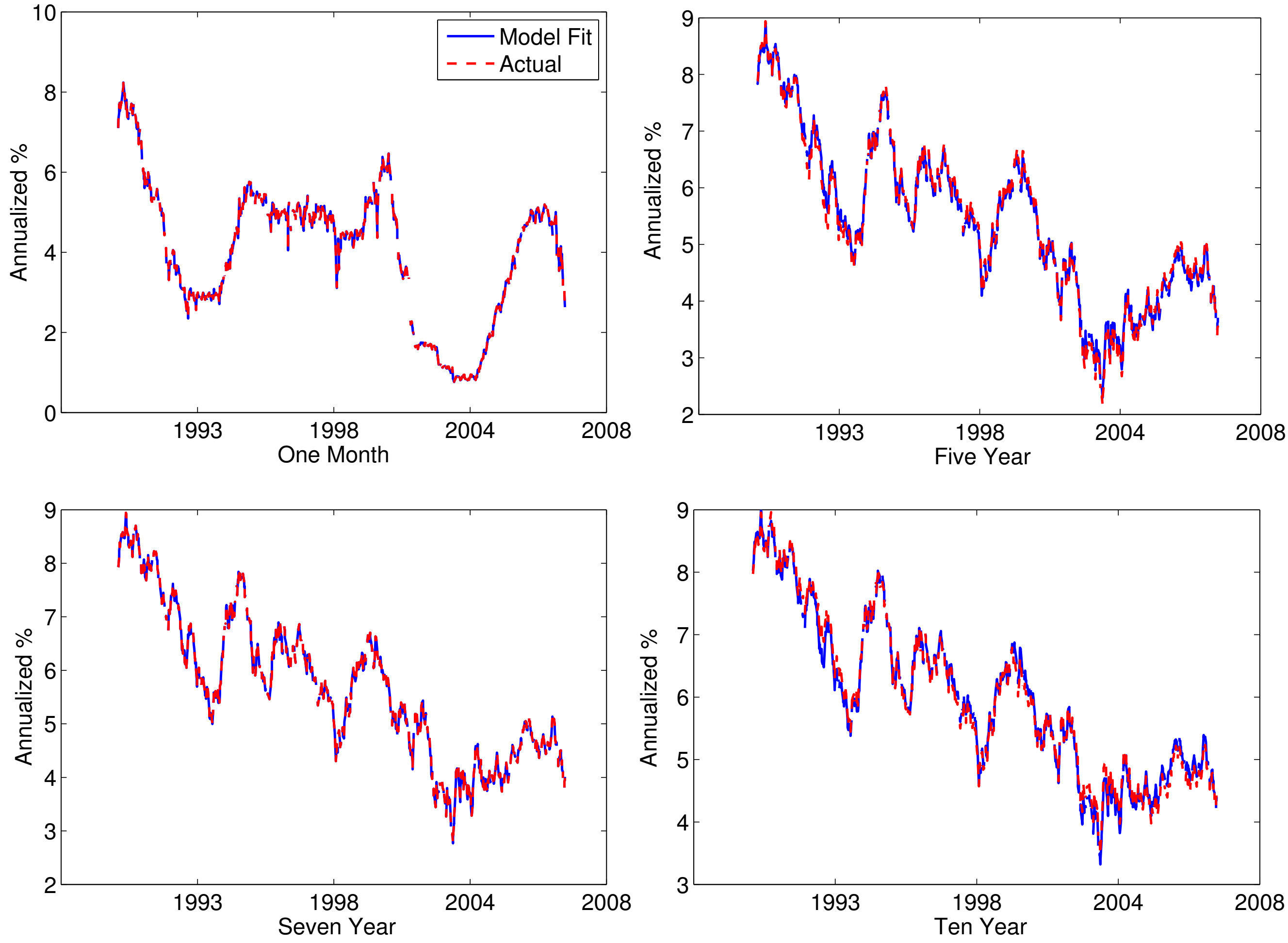

Figure 3: Model fit vs. Data for the Bond Yields

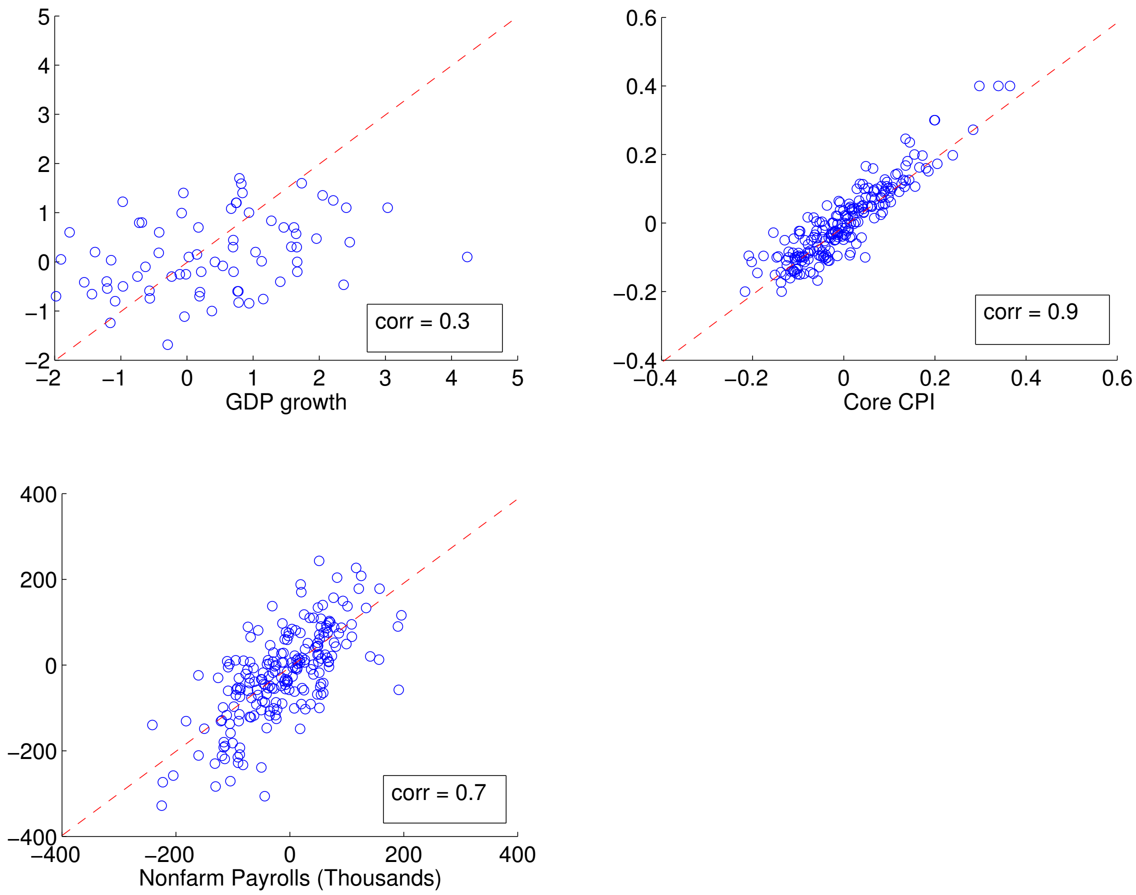

Figure 4: Model Implied and Actual Surprises.

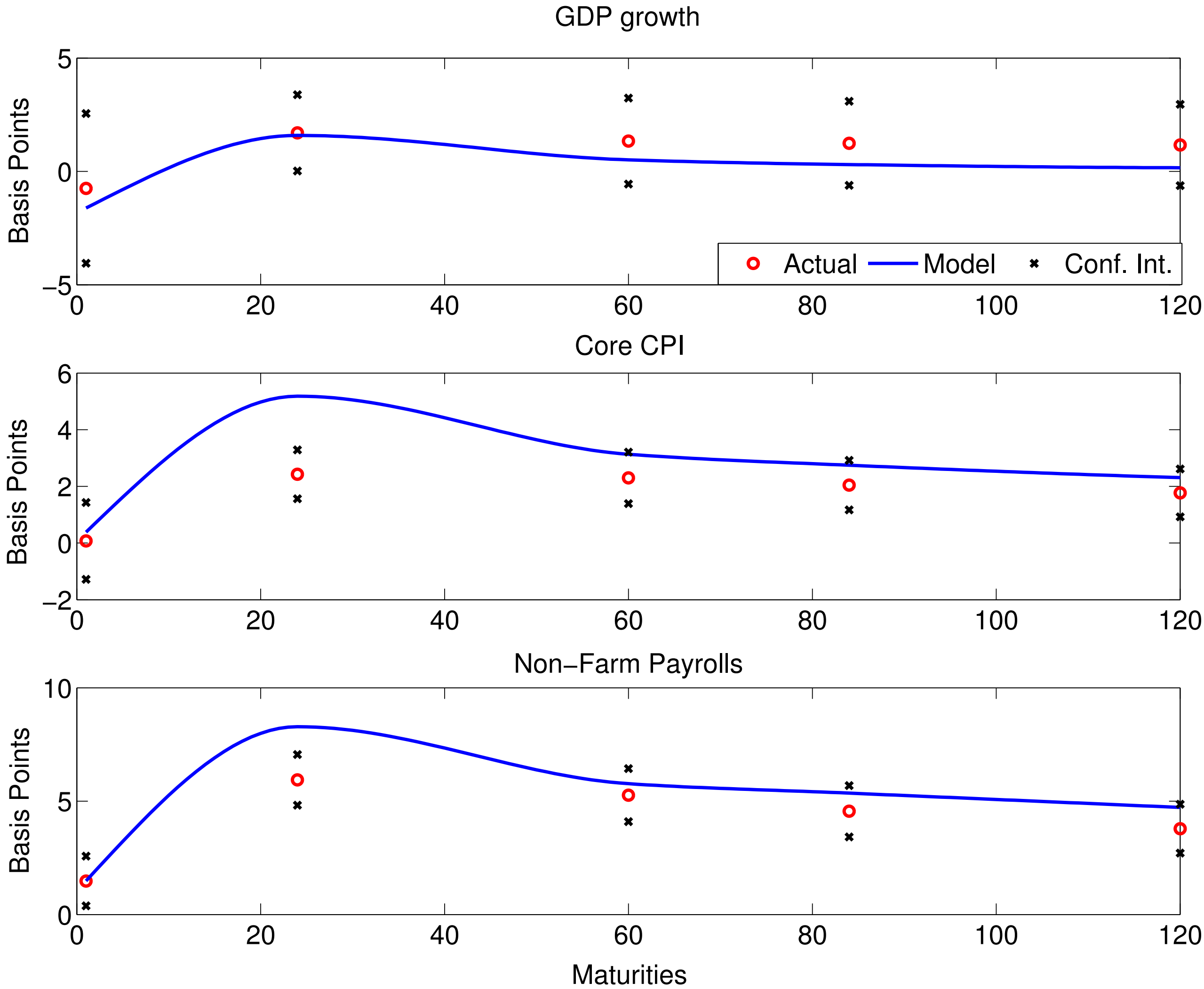

Figure 5: Event-study responses vs. Model Responses

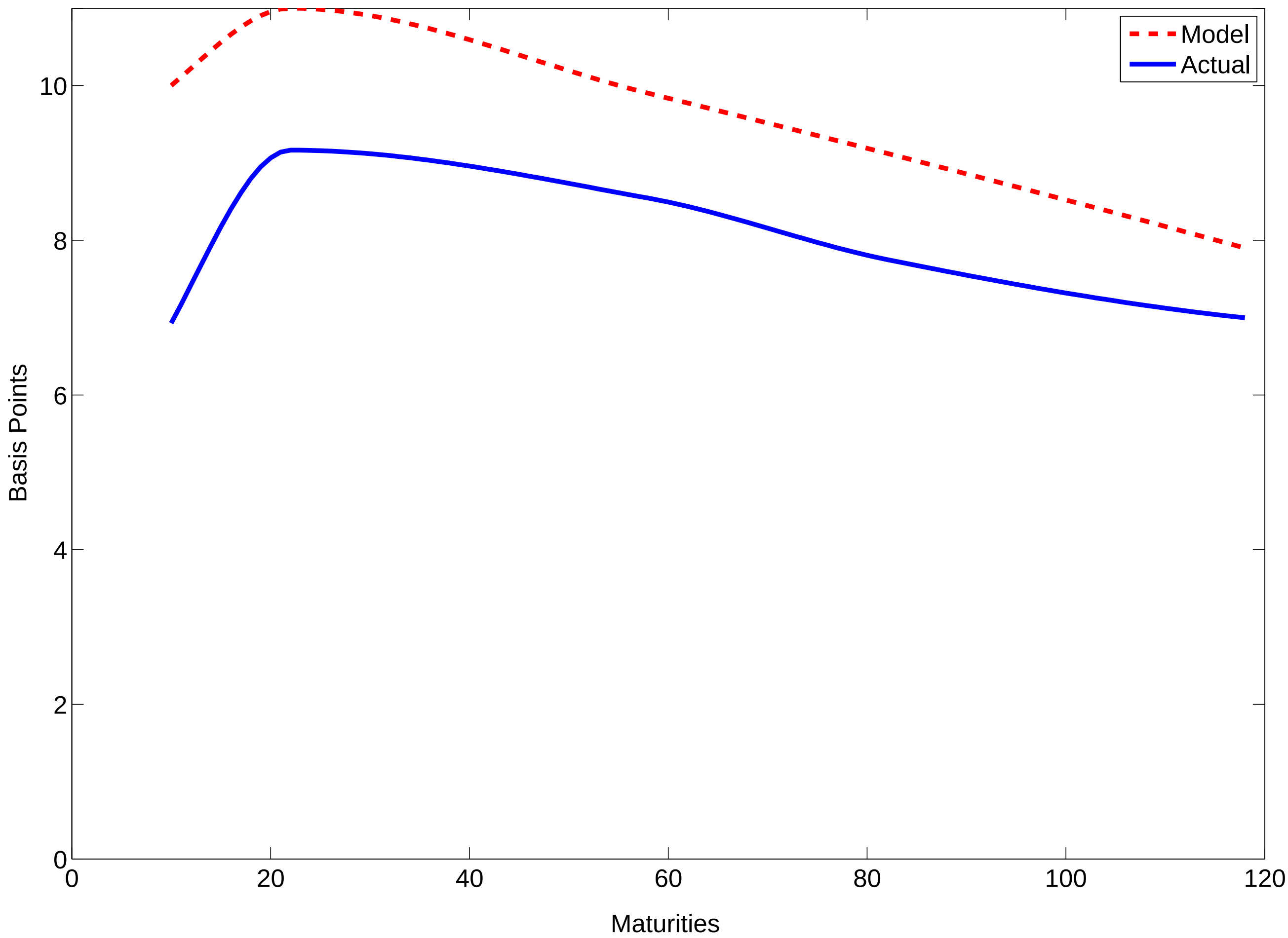

Figure 6: Actual vs. Model Volatility Curves

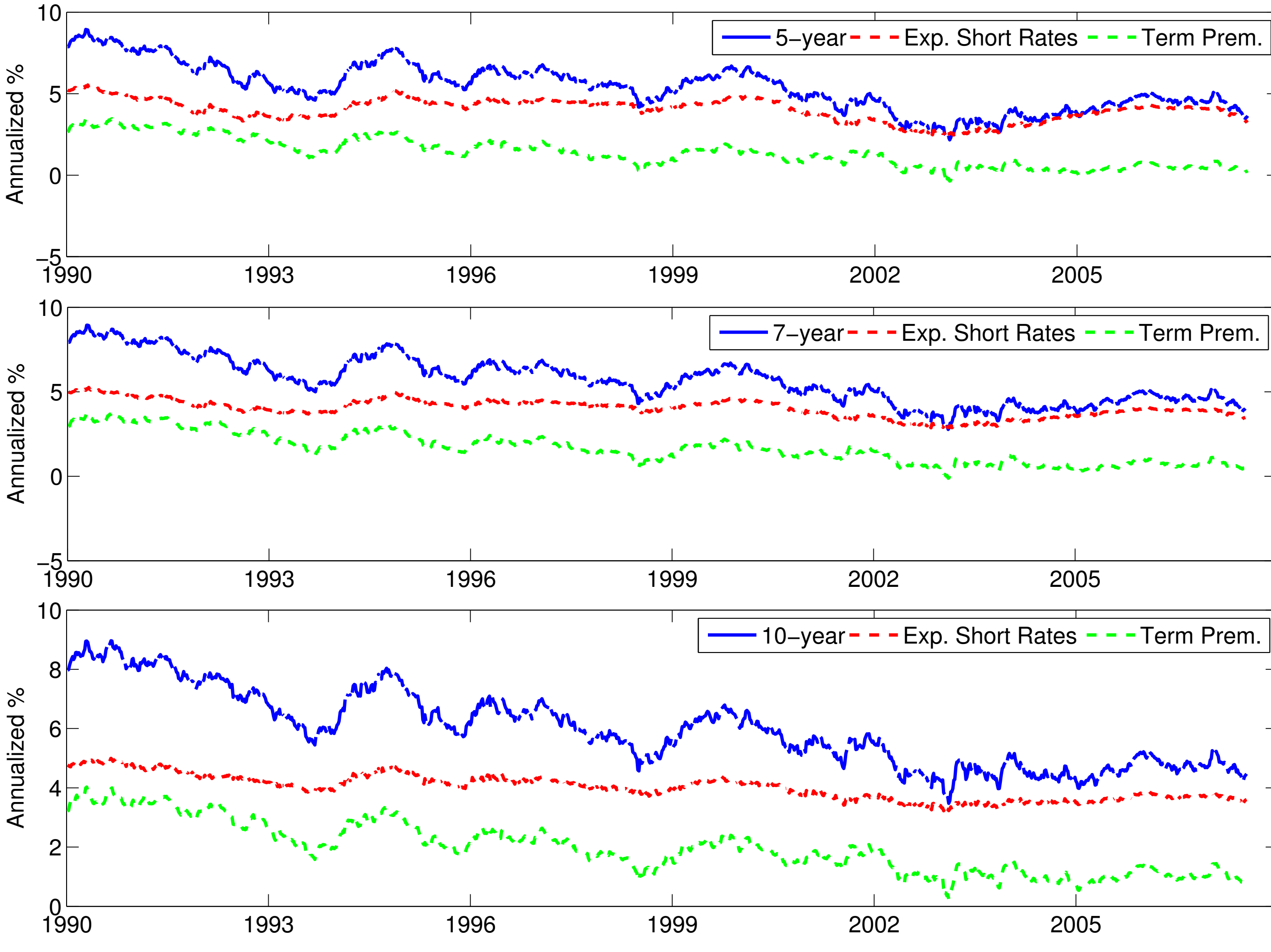

Figure 7: Decomposition of 5-year, 7-year and 10-year nominal zero coupon bond yields

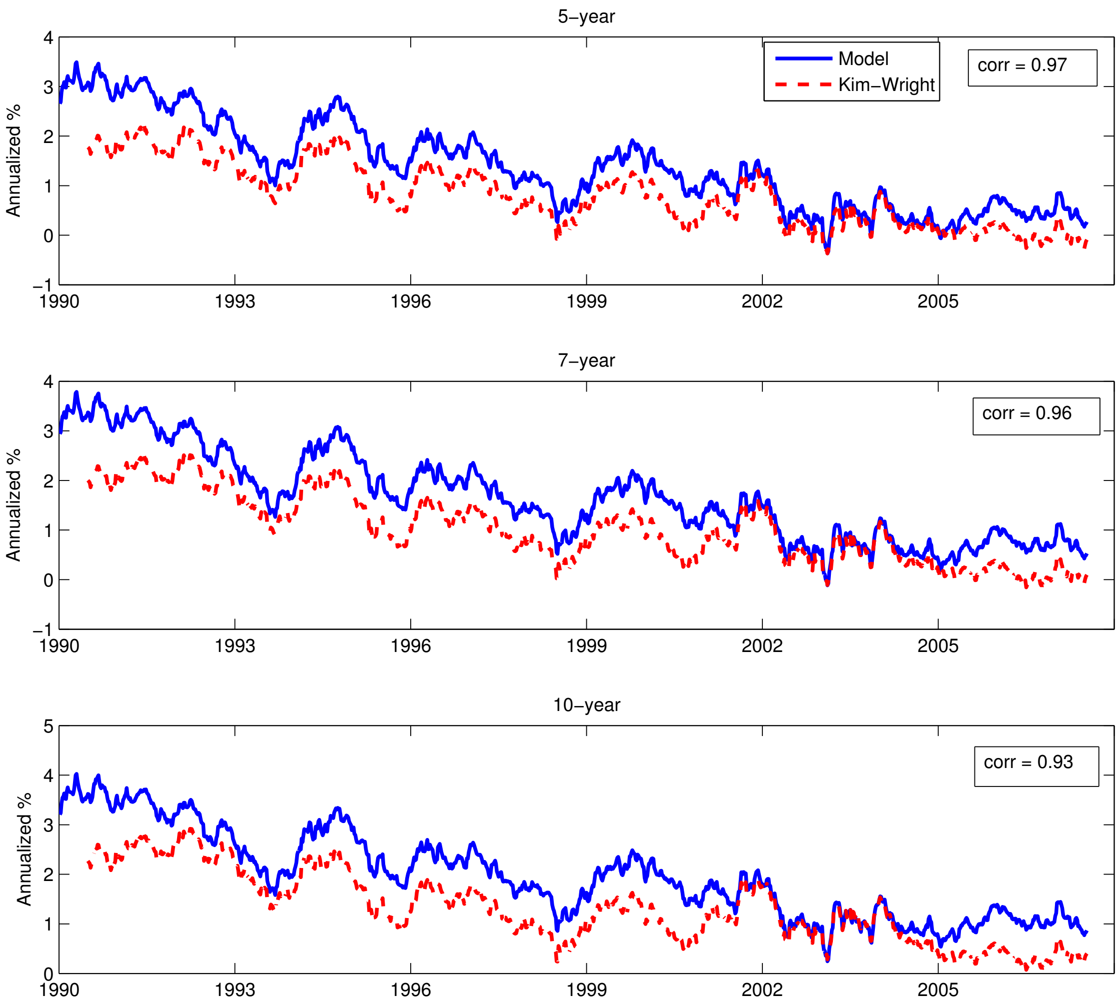

Figure 8: Kim-Wright and Model Implied 5-year, 7-year and 10-year Term Premia.

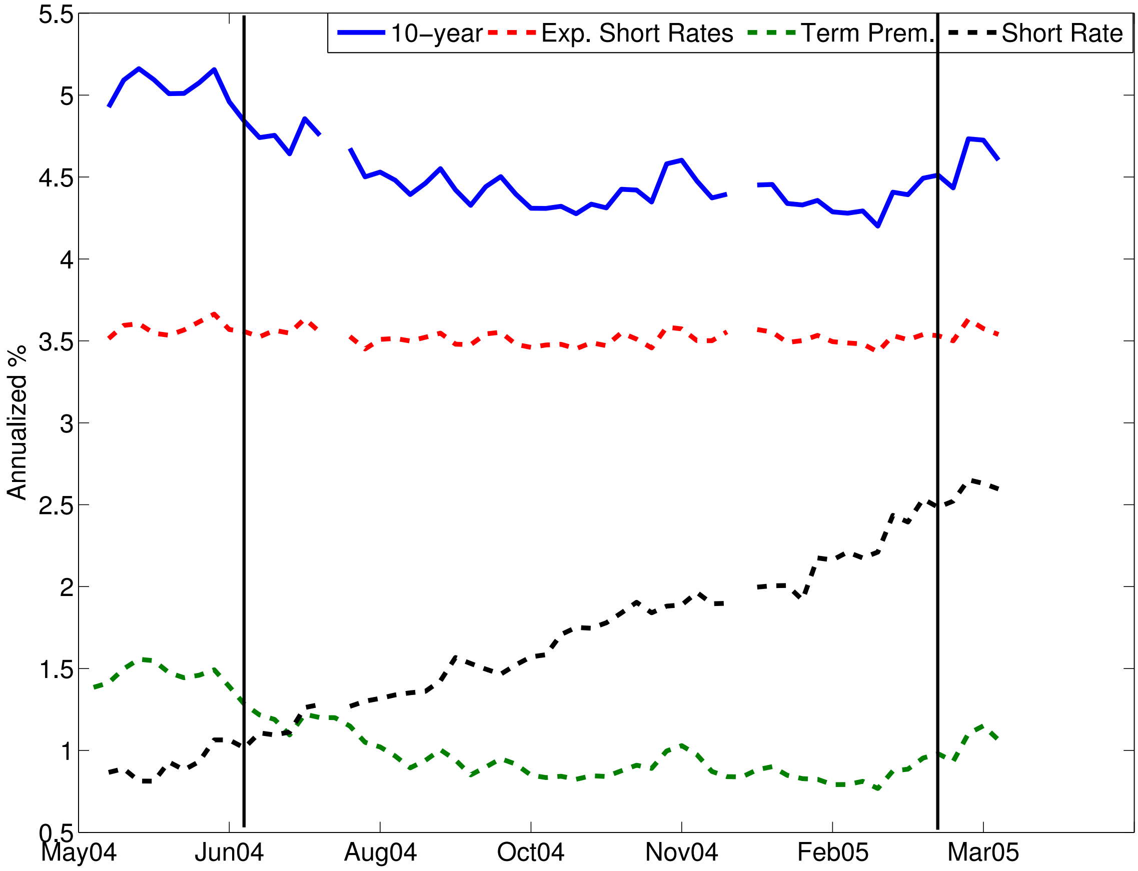

Figure 9: Behavior of the 10-year nominal zero coupon bond yields during the conundrum period. The conundrum period is denoted by two vertical lines on the graph which corresponds to June 29, 2004 and February 2, 2005.

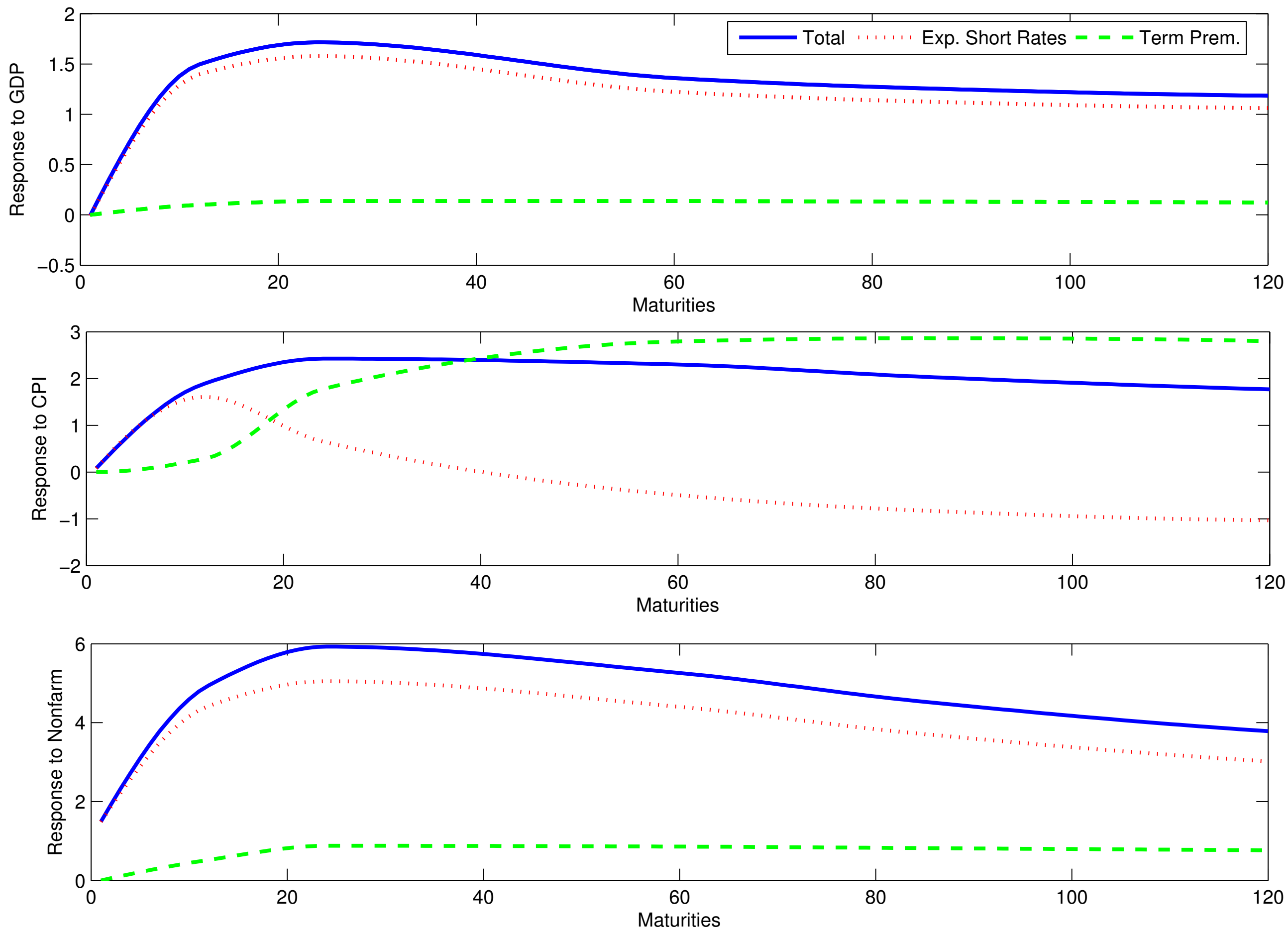

Figure 10: Responses of bond yields and their components to announcements.

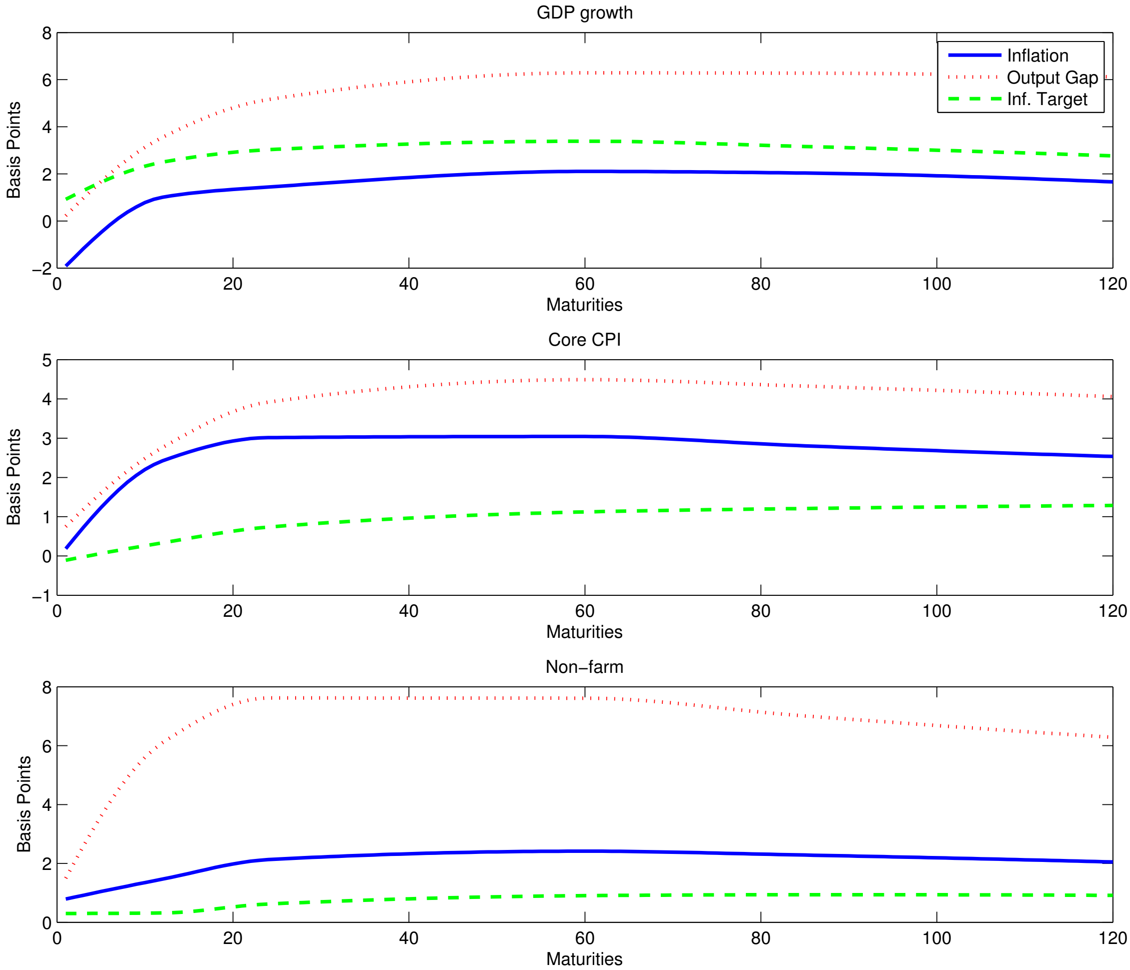

Figure 11: Responses of bond yields to state variables

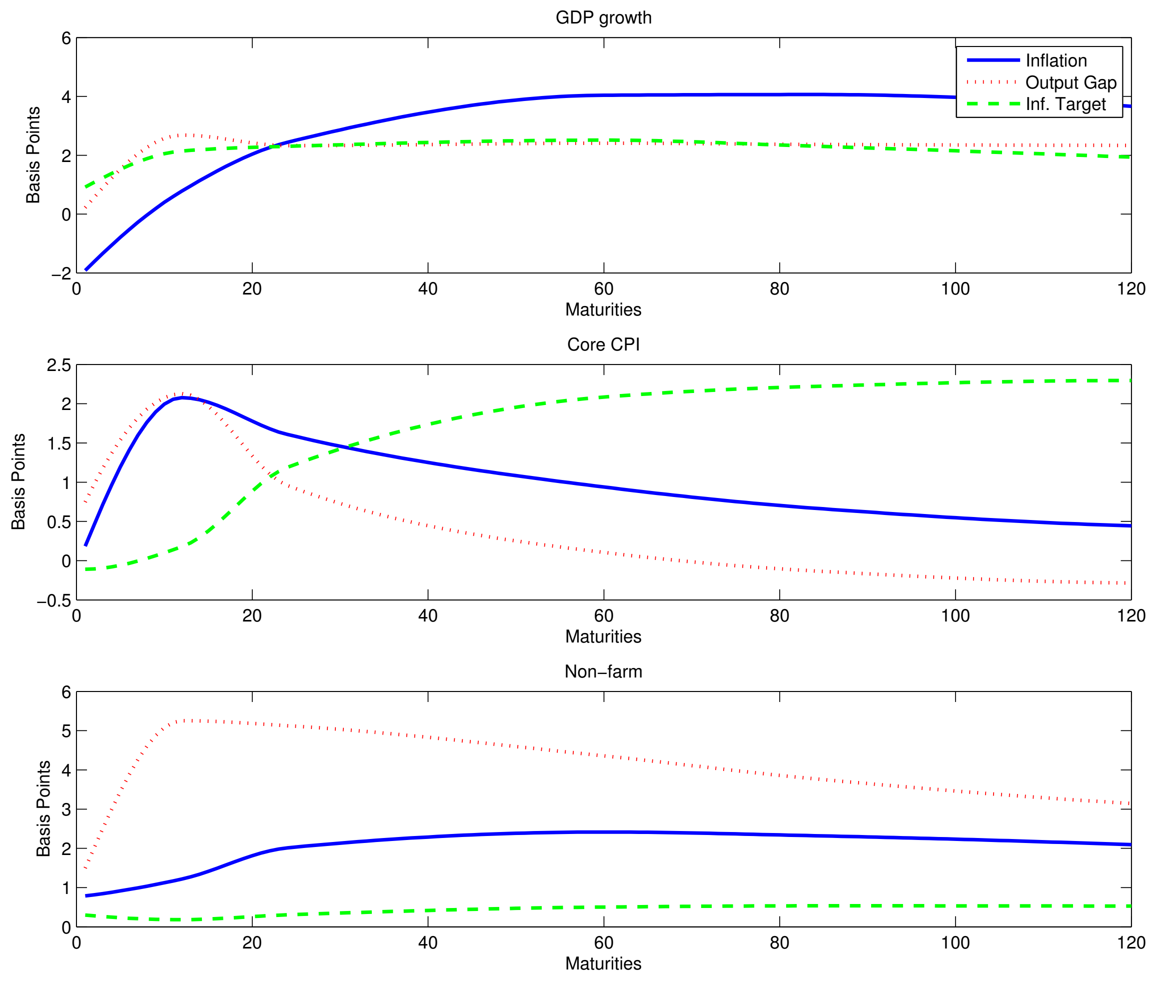

Figure 12: Responses of average expected short rates to state variables

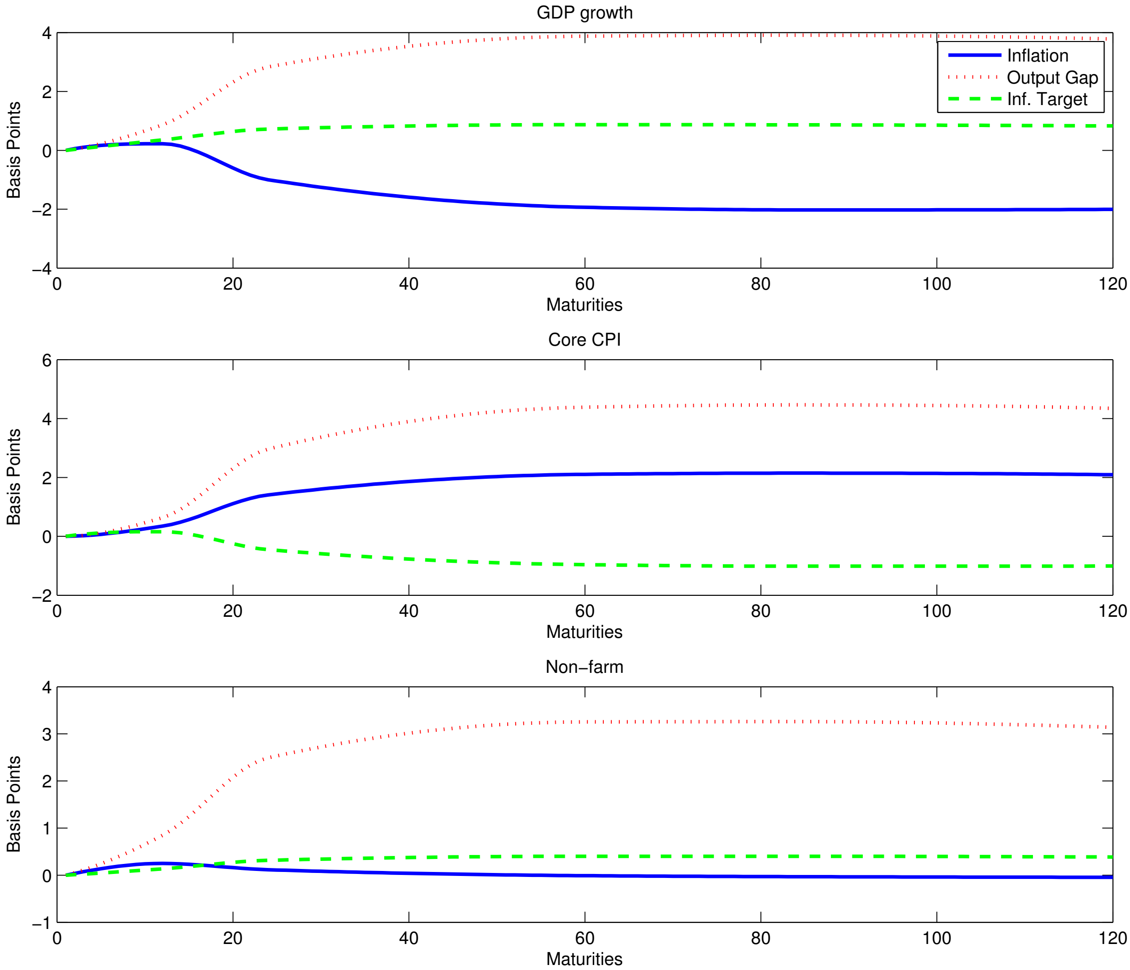

Figure 13: Responses of term premia to state variables

	Prior mode	Prior Std. Dev.	Dist.	Post. mean	Post. Std. Dev.
γ	2	0.1	Normal	1.64	0.0353
φ	5	1	Normal	5.14	0.0217
η	1	0.05	Normal	1.01	0.0234
θ	0.9	0.1	Beta	0.94	0.0558
ω	0.3	0.05	Beta	0.17	0.0247
β	0.99	0.03	Beta	0.98	0.0135
ρ	0.5	0.15	Beta	0.51	0.0349
λ_y	0.5	0.1	Beta	0.46	0.0083
λ_i	0.5	0.1	Beta	0.54	0.0137
ρ_{π^*}	0.95	0.01	Beta	0.93	0.0394
ϕ_1	-	-	Uniform(0,∞)	1.02	0.06
ϕ_2	-	-	Uniform(0,∞)	0.94	0.05
ρ_b	0.95	0.01	Beta	0.95	0.0366
σ_a	-	-	Uniform(0,∞)	0.21	0.096
σ_y	-	-	Uniform(0,∞)	0.12	0.0928
σ_π	-	-	Uniform(0,∞)	0.29	0.01
σ_{π^*}	0.01	0.01	Inverse Gamma	0.4	0.04
σ_b	0.01	0.01	Inverse Gamma	0.09	0.038
$\sigma_y^{v,cb} \times 12$	-	-	Uniform(0,σ_y)	0.8	0.02
$\sigma_\pi^{v,cb}$	-	-	Uniform(0,σ_π)	0.03	0.003
$\sigma_n^{v,cb} \times 1000$	-	-	Uniform(0,σ_n)	72	3.2
$\sigma_{\pi^*}^{v,cb} \times 12$	0.013	0.1	Inverse Gamma	0.3	0.0032
$\sigma_y^v \times 12$	-	-	Uniform(0,σ_y)	1.6	0.08
σ_π^v	-	-	Uniform(0,σ_π)	0.05	0.0032
$\sigma_n^v \times 1000$	-	-	Uniform(0,σ_n)	132	6.1
$\sigma_{\pi^*}^v \times 12$	0.013	0.1	Inverse Gamma	0.44	0.01
$\sigma_{y,survey}^v \times 12$	-	-	Uniform(0,σ_y)	1.09	0.0099
$\sigma_{\pi,survey}^v$	-	-	Uniform(0,σ_π)	0.01	0.007
$\sigma_{n,survey}^v \times 1000$	-	-	Uniform(0,σ_n)	28	3
$\sigma_{\mathcal{Y}^{1m}}^v \times 1200$	10 bps	0.1	Normal	10 bps	0.46
$\sigma_{\mathcal{Y}^{1y}}^v \times 1200$	10 bps	0.1	Normal	7 bps	0.22
$\sigma_{\mathcal{Y}^{2y}}^v \times 1200$	10 bps	0.1	Normal	6 bps	0.3
$\sigma_{\mathcal{Y}^{5y}}^v \times 1200$	10 bps	0.1	Normal	8 bps	0.02
$\sigma_{\mathcal{Y}^{7y}}^v \times 1200$	10 bps	0.1	Normal	1 bps	0.0046
$\sigma_{\mathcal{Y}^{10y}}^v \times 1200$	10 bps	0.1	Normal	9 bps	0.03

Table 1: Priors and Posterior estimates of the model parameters.

Notes: Posterior estimates based on announcement/survey data of the real GDP growth, core CPI inflation, non-farm payrolls, Blue Chip inflation expectations from and bond yields of one month, one year, two year, five year, seven year and ten year maturities from January 1990 to December 2007. Posterior parameter standard deviations were calculated using 300,000 draws from the random walk Metropolis-Hastings algorithm using the first 50,000 draws as burn-in.

	std(Δ5-year)	std(ΔExp)	std(ΔTP)	corr(ΔExp , ΔTP)
GDP	0.0794	0.05	0.05	0.14
Core CPI	0.0707	0.045	0.05	0.03
Non-farm	0.1	0.05	0.05	0.45

Table 2: Variance decomposition of five year zero coupon yields around announcements. Table shows the standard deviations, denoted as *std*(.), of the changes in five year yield, average expected short rates in five years and the term premium.

	std(Δ7-year)	std(ΔExp)	std(ΔTP)	corr(ΔExp , ΔTP)
GDP	0.0775	0.048	0.055	0.1
Core CPI	0.0671	0.04	0.05	-0.02
Non-farm	0.0938	0.05	0.06	0.49

Table 3: Variance decomposition of seven year zero coupon yields around announcements. Table shows the standard deviations, denoted as *std*(.), of the changes in seven year yield, average expected short rates in seven years and the term premium.

	std(Δ10-year)	std(ΔExp)	std(ΔTP)	corr(ΔExp , ΔTP)
GDP	0.0748	0.05	0.048	0.12
Core CPI	0.0648	0.04	0.05	-0.07
Non-farm	0.0866	0.05	0.049	0.51

Table 4: Variance decomposition of ten year zero coupon yields around announcements. Table shows the standard deviations, denoted as *std*(.), of the changes in ten year yield, average expected short rates in seven years and the term premium.