The Effects of Treasury Debt Supply on Macroeconomic and Term Structure Dynamics

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Abstract

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1 Introduction

The relationship between the macroeconomy and asset pricing has been a long-standing area of research. One aspect that has received much attention is the interaction between the macroeconomy and the term structure of interest rates, and how this interaction is shaped by the conduct of monetary policy. Whereas the monetary policy rule determines the comovement of macroeconomic aggregates and expectations of future short-term interest rates, its implications, and those of other aspects of the macroeconomy, for the determination of term premia are less well understood. The use of large-scale asset purchases or "quantitative easing" (QE) by a number of central banks since the intensification of the financial crisis in 2008 has lent urgency to a better understanding of macroeconomic determinants of the yield curve, as those policy measures are thought to operate to a large extent by influencing term premia.

In this paper, we provide empirical evidence on the joint response of US Treasury yields, the embedded term premia, and macroeconomic variables to changes in the supply of Treasury securities. To do so, we follow the approach first developed in Ang and Piazzesi (2003) of combining an affine model of the term structure of Treasury yields with the assumption that the factors that matter for pricing these bonds are macroeconomic variables whose joint dynamics can be described by a linear VAR model. The focus of our analysis is on identifying exogenous variations in Treasury supply emanating from fiscal and monetary policy and to estimate the response of yields and term premia across the maturity spectrum. The objective is to provide empirical evidence on these responses that is based on minimal identifying assumptions, so as to uncover "stylized facts" for future study in more fully specified structural models.

An important motivation, as mentioned before, is the recent use of QE by a number of central banks. The macroeconomic effects of these policies are widely debated (some relevant studies will be discussed below). In many instances, the study of these effects has proceeded in two steps. First, there are a number of case studies that document the response of various asset prices to QE-related news in narrow event windows. Whereas with the proper choice of event window one can hope to capture the asset price responses of such news, the responses of slower-moving macroeconomic variables such as output and inflation can of course not be measured in this way. Therefore, in a second step, the asset price responses identified from the event studies are used in structural macroeconomic models to obtain results for the effects of QE on output, the unemployment rate, and inflation. This second step involves the choice of a particular structural model which implies numerous, and often contentious, assumptions about the transmission mechanism of the financial effects of QE.

In contrast to this two-step approach to the estimation of QE effects, in this study we model the dynamics of macro and financial variables jointly. In the spirit of the structural VAR literature, we aim to reduce the imposition of prior assumptions about the dynamic responses of macroeconomic variables to QE events to a minimum of identifying assumptions. We make use of data on both total supply of marketable securities by the Treasury and Treasury holdings by the Federal Reserve and foreign official institutions, and build on the literature on fiscal SVARs to disentangle exogenous innovations to Treasury supply emanating from fiscal policy from those emanating from monetary policy. Doing so is arguably important, as the effects of the former reflect both the effects of changes in the amount of Treasuries held in private portfolios and the effects of the tax and spending decisions that bring about the change in Treasury supply, whereas QE-style changes in Treasury supply are not associated with fiscal policy changes.

Aside from the study of the dynamic effects of QE, there is a long-standing literature examining the information content in the term structure, and especially its slope, for future economic activity. A premise of the macro-finance literature is that risk premia, including the term premia for bearing duration risk, are endogenous variables that can be affected by various macroeconomic disturbances. Hence, reduced-form regressions of measures of economic activity on the slope or some other measure of term premia are unlikely to uncover the partial effect of a change in term premia induced by a policy intervention such as QE. The impulse response functions of yields and term premia to various shocks produced by our model shed some light on the correlation patterns between term premia and real activity induced by these shocks.

Our [highly preliminary] findings are that (i) exogenous increases in Federal Reserve or foreign official holdings raise output significantly and by similar amounts as estimated in previous studies, but that inflation rises substantially, and that these effects are tempered by an increase in the short-term interest rate; (ii) that exogenous increases in Treasury supply of the same size induced by fiscal policy lead to significantly larger responses of output, inflation, and short-term interest rates; and (iii) that long-term Treasury yields rise in response to both types of shocks, but that the response to an increase in Federal Reserve or foreign official holdings is associated with a sizeable and significant decline in term premia, whereas fiscal policy-induced Treasury supply changes are not.

In the remainder of this introduction we discuss a few studies that are closely related to ours. In section 2, we discuss the specification of our term structure model, the identification strategy, and the data we use. Section 3 presents our empirical findings, and section 4 offers conclusions. The appendix spells out further details regarding model specification, data, and identification.

1.1 Related literature

The studies by Ang and Piazzesi (2003) and by Ang, Piazzesi, and Wei (2006) pioneered the use of macroeconomic variables as factors in affine term structure models, based on the intuition that, if the central bank varies short-term interest rates systematically in response to economic variables, these variables must be relevant for bond pricing. However, this use of macroeconomic variables as pricing factors is not uncontroversial. If macro variables were the only factors affecting bond pricing, regressions of macro variables on yields should show high R^2s, whereas in the data the R^2 is very small, especially for real growth variables.[1] Instead, there seems to be information in macro variables that is relevant for predicting future short-term rates and future excess returns on longer-term bonds, but this information does not affect current bond pricing, a phenomenon termed "unspanned macro risks" by Joslin, Priebsch, and Singleton (2014). In this paper, we follow their approach and model the cross section of yields as driven by yield factors only whereas the expected future yields and term premiums are driven by current yield as well as macro and supply variables.

Our study is also related to the rapidly growing literature on the supply and demand effects on the government bond market. A number of papers documented that Treasury yields and future returns to Treasury securities tend to be lower when the debt-to-GDP ratio is lower[2] or when there are more demand for Treasury securities from foreign investors[3].

[1](See Orphanides and Wei (2012), Kim (2009), Duffee (2011), Gürkaynak and Wright (2012) and Duffee (2013) for more discussions).

[2]See, for example, Greenwood and Vayanos (2010, 2014), Hamilton and Wu (2012), Krishnamurthy and Vissing-Jorgensen (2012), Laubach (2009), and Thornton (2012).

[3]See Warnock and Warnock (2009), Beltran, Kretchmer, Marquez, and Thomas (2013), Kaminska, Vayanos, and Zinna (2011), Kaminska and Zinna (2014), Jaramillo and Zhang (2013)

More recently, various central banks' unconventional monetary policy responses to the 2007-8 financial crisis provided additional data for studying the link between bond supply and bond yields. Early case studies of the responses of long-term (Treasury and private) yields to QE-related announcements include Gagnon, Raskin, Remache, and Sack (2011) and Krishnamurthy and Vissing-Jorgensen (2011) for the U.S. and Joyce, Tong, and Woods (2011) and McLaren, Banerjee, and Latto (2014) for the U.K. All studies find significant effects of announcements that raised expectations of central bank purchases of Treasuries or agency MBS on yields of Treasury securities, MBS, and to a lesser extent corporates. Li and Wei (2013) obtain comparably estimates using a term structure model. Two earlier episodes in the U.S., the Operation Twist and the Treasury bond buy-back program, that similarly reduced the duration supply of Treasury securities are also found to have modestly reduced bond yields at the time.[4] An important channel through which a lower supply of Treasury securities or a higher demand from price-insensitive investors such as central banks or foreign investors reduce bond yields is by reducing the risk compensation price-sensitive investors demand for bearing interest rate risks–the "duration channel" as proposed in Vayanos and Vila (2009).[5]

(To be completed.)

Kiley (2013b,a).

Macroeconomic effects of unconventional monetary policy: Chung, Laforte, Reifschneider, and Williams (2012), Chen, Cúrdia, and Ferrero (2011).

Li and Wei (2013)

Macroeconomic effects of fiscal policy: Blanchard and Perotti (2002), Dai and Philippon (2004).

2 A First Look at the Data

In this section we conduct some exploratory analysis of the data. This will also serve as the motivation for the more structural term structure model introduced in the following

[4]See Swanson (2011) for the former and Bernanke, Reinhart, and Sack (2004) and Greenwood and Vayanos (2010) for the latter.

[5]Other channels have been proposed in the literature, including the local supply channel by D'Amico and King (2010), the "signaling channel" by Bauer and Rudebusch (2012), and the "safety premium" channel by Krishnamurthy and Vissing-Jorgensen (2011). Krishnamurthy and Vissing-Jorgensen (2011) and D'Amico, English, López-Salido, and Nelson (2012) provide evidence on the relative strength of various channels.

Table 1: Regressions of 10-Year Treasury Yields and Term Premiums

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	10-year yields				10-year term premiums			
short rate	0.69	0.69	0.63	0.57	0.26	0.25	0.19	0.15
	(0.03)	(0.05)	(0.04)	(0.06)	(0.02)	0.04)	(0.03)	(0.04)
expected growth		0.18		0.16		0.14		0.13
		(0.07)		(0.07)		(0.05)		(0.05)
expected inflation		0.10		0.14		0.05		0.08
		(0.08)		(0.08)		(0.06)		(0.06)
CP factor								
total supply			0.01	0.00			0.01	0.01
			(0.02)	(0.02)			(0.01)	(0.01)
official purchases			-0.10	-0.10			-0.08	-0.08
			(0.03)	(0.03)			(0.02)	(0.02)
Adjusted R^2	0.80	0.83	0.84	0.85	0.55	0.57	0.60	0.62

This paper reports regression results of 10-year Treasury yields and Kim and Wright (2005) 10-year term premium estimates on the short rate, survey forecasts of next-quarter real GDP growth and inflation, and Treasury supply variables. T-statistics are reported in parentheses.

sections.

Table 1 regresses the 10-year Treasury yield and the Kim and Wright (2005) measures of the 10-year term premium on combinations of 1-quarter short rate, Blue Chip survey forecasts of next-quarter GDP growth and next-quarter CPI inflation, total Treasury debt-to-GDP ratio, and foreign official and Federal Reserve holdings of Treasury securities. Consistent with previous findings, we find that an increase in the official Treasury holdings is typically accompanied by a decline in Treasury yields and term premiums,

3 The term structure model with supply factors

In this section we specify an affine model of the term structure of nominal US Treasury zero-coupon securities with macroeconomic and supply variables. Our choice of factors reflects views on which fundamental sources of economic uncertainty are likely reflected in bond prices. These include long-run risks, such as perceptions of persistent changes in inflation or the equilibrium real interest rate, as well as variables relevant for determining the amount of duration risk held by private investors, such as total Treasury supply and foreign and domestic official holdings of Treasuries. We allow the macro and supply factors to be unspanned by the cross section of yields following Joslin, Priebsch, and Singleton (2014), motivated by empirical evidence that yield changes seem to explain only a small portion of the variations in those factors.

3.1 An arbitrage-free term structure model with unspanned macro and supply factors

The model used to estimate the macroeconomic and term structure effects of Treasury supply shocks consists of a description of the relationships between major macroeconomic and Treasury supply variables, and a specification of the stochastic discount factor that ensures that pricing of bonds at various maturities is arbitrage-free. The model prices nominal zero-coupon bonds that are free of default risk.

Following Joslin, Priebsch, and Singleton (2014), we assume that the current levels of yields are completely determined by yield factors alone. This is achieved by assuming that those yield factors follow an autonomous process under the risk-neutral measure $\mathbb{Q}$ and that the short rate loads on those factors only:

$$\mathcal{P}_t = \mu_{\mathcal{P}}^{\mathbb{Q}} + \Phi_{\mathcal{P}}^{\mathbb{Q}} \mathcal{P}_{t-1} + \Sigma_{\mathcal{P}}^{\mathbb{Q}} \epsilon_{\mathcal{P},t}^{\mathbb{Q}}, \ t = 1, \ldots, T, \tag{1}$$

$$y_t^1 = \delta_0 + \delta_1' \mathcal{P}_t \tag{2}$$

We denote the yield factors by $\mathcal{P}_t = \{\mathcal{P}_t^1, \mathcal{P}_t^2, \mathcal{P}_t^3\}$, and measure them using either the first three principal component factors (PCs) of yields or the level, slope, curvature of the yield curve, defined as the 1-quarter yield, the difference between the 10-year and the one-quarter yield, and the sum of 1-quarter and 10-year yields minus two times the 2-year yield, respectively. Either set of measures explain more than 99% of the time variations of yields at all maturities.

We further adopt the Joslin, Singleton, and Zhu (2011) canonical form, under which the $\mathbb{Q}$ distribution of $\mathcal{P}_t$ is fully characterized by the parameters $\Theta_{\mathcal{P}}^{\mathbb{Q}} \equiv (\kappa_\infty^{\mathbb{Q}}, \lambda^{\mathbb{Q}}, \Sigma_{\mathcal{P}})$, and $\mu^{\mathbb{Q}}$, $\Phi^{\mathbb{Q}}$, δ_0, and δ_1 are all explicit functions of $\Theta_{\mathcal{P}}^{\mathbb{Q}}$. This model then implies that the yield on a nominal zero-coupon bond with n periods to maturity is a linear function of the yield factors:

$$y_{n,t} = a_n(\Theta^{\mathbb{Q}}_{\mathcal{P}}) + b_n(\Theta^{\mathbb{Q}}_{\mathcal{P}})'\mathcal{P}_t, \tag{3}$$

where the coefficients $a_n,\ b_n$ are determined recursively.

Under the physical measure $\mathbb{P}$, however, those yield factors form part of a larger-scale VAR that also includes macroeconomic factors.

$$X_t = \mu^{\mathbb{P}} + \Phi^{\mathbb{P}} X_{t-1} + \Sigma^{\mathbb{P}} \epsilon^{\mathbb{P}}_t,\ t = 1, \ldots, T, \tag{4}$$

As standard in the macro term structure literature, we include a real growth measure and an inflation measure. For reasons that will be discussed below, we consider inflation π_t and the one-period real interest rate r_t as each containing a time-varying asymptote, represented by an overhead bar, and denote the stationary deviation from this asymptote or trend by a tilde:

$$\pi_t = \bar{\pi}_t + \tilde{\pi}_t \tag{5}$$

$$r_t = \bar{r}_t + \tilde{r}_t \tag{6}$$

The trend component of inflation may reflect time-varying perceptions of the central bank's inflation objective. Since we assume that $\tilde{\pi}_t$ has an unconditional mean of zero, the infinite-horizon expectation of inflation at time t is given by $\bar{\pi}_t$. Analogously, the trend component of the one-period real interest rate may reflect time-varying perceptions of trend productivity growth or changes in risk attitudes that affect the infinite-horizon expectation of the equilibrium real rate of return on short-term risk-free assets.[6]

We detrend the first yield factor using the inflation and real rate trends. For example, the detrended PC1 factor, $\tilde{\mathcal{P}}^1_t$, is linked to the observed PC1 factor $\mathcal{P}^1_t$ by

$$\mathcal{P}^1_t = \tilde{\mathcal{P}}^1_t + \bar{\pi}_t + \bar{r}_t, \tag{7}$$

and the detrended one-period nominal yield, $\tilde{y}^1_t$, is defined similarly. With this notation, we define the vector of state variables or factors driving yields as

$$x_t = [\bar{\pi}_t,\ \bar{r}_t,\ \tilde{\pi}_t,\ q_t,\ t_t,\ s_t,\ \tilde{\mathcal{P}}^1_t,\ \mathcal{P}^2_t,\ \mathcal{P}^3_t]' \tag{8}$$

where $q_t,\ t_t$, and s_t are measures of real activity, total marketable Treasury debt, and domestic and foreign official holdings of Treasury securities, respectively.

[6]Spencer (2008) presents a term structure model in which yields also depend on time-varying asymptotes of inflation and the real short rate.

The state vector x_t is assumed to follow a VAR(q) under the physical measure:

$$x_t = \mu^{\mathbb{P}} + \phi_1^{\mathbb{P}} x_{t-1} + \ldots + \phi_q^{\mathbb{P}} x_{t-q} + \Sigma^{\mathbb{P}} \epsilon_t^{\mathbb{P}}, \ t = 1, \ldots, T \tag{9}$$

Under the assumption (discussed further below) that trend inflation and the trend real short-term interest rate follow univariate random walks, whereas the other elements of x_t follow a VAR(q), the the vectors and matrices in (4) are of the form

$$X_t \equiv \begin{bmatrix} x_t \\ x_{t-1} \\ \vdots \\ x_{t-q} \end{bmatrix}, \ \epsilon_t \equiv \begin{bmatrix} \epsilon_t \\ 0 \\ \vdots \\ 0 \end{bmatrix}, \ \Phi \equiv \begin{bmatrix} I_2 & 0 & \ldots & 0 & 0 \\ 0 & \phi_1 & \ldots & \phi_q & 0 \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & \ldots & 0 & 0 \\ 0 & 0 & \ldots & 0 & 0 \end{bmatrix}$$

where $\phi_i, \ i = 1 \ldots q$ are coefficient matrices of size 7×7, and I_n symbolizes the identity matrix of size $n \times n$.

Since the nominal short rate only loads on the yield factors, only those factors are priced in the nominal Treasury market:

$$\log M_{t+1} = -y_t^1 - \frac{1}{2}\Lambda_t^{'}\Lambda_t - \Lambda_t^{'}\epsilon_{t+1}^{\mathbb{Q}}, \tag{10}$$

where the price of risk parameters is determined by the parameters governing the $\mathbb{P}$- and the $\mathbb{Q}$-VARs:

$$\lambda_t = \lambda_0 + \lambda_1 \mathcal{P}_t \tag{11}$$

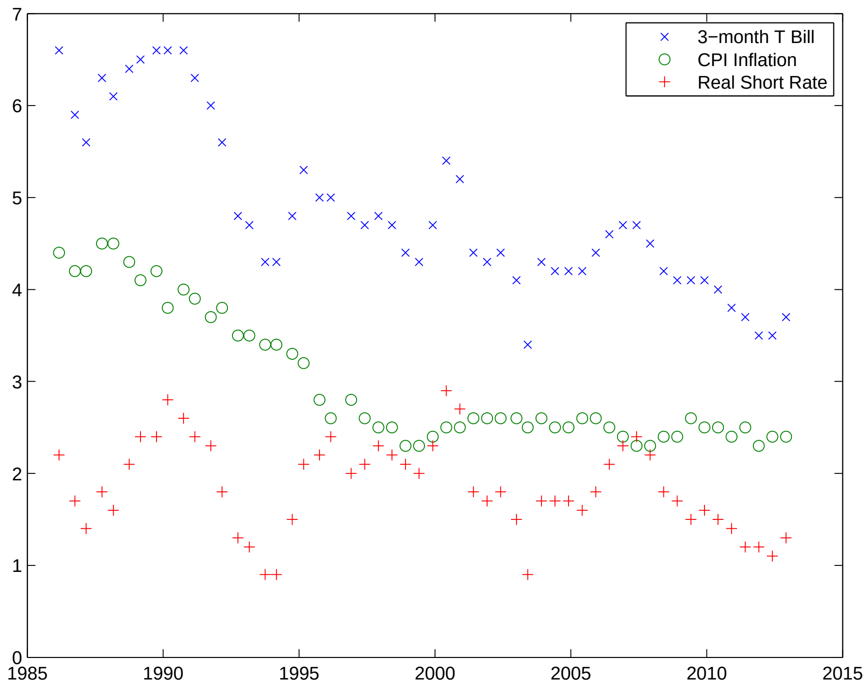
More details are provided in Appendix A. This model therefore falls into the category of essentially-affine term structure models (e.g. Duffee (2002)).

3.2 Data and survey expectations used in estimation

For our empirical implementation we specify the model at quarterly frequency. This choice reflects that fact that we are interested in measuring the effects of changes in Treasury supply stemming from fiscal or monetary policy actions on macroeconomic variables and yields jointly, and recognizes that a lot variations in yields at higher data frequencies can likely not be attributed to these factors.

Because the decomposition of yields into contributions from expected future short-term interest rates and term premia depends on an accurate modeling of financial market participants' expectations, we follow Kim and Orphanides (2012) by making extensive use of

Figure 1: Long-Horizon Expectations of 3-month Yields and Inflation, Blue Chip



survey expectations in both estimation and model evaluation. In particular, Kim and Orphanides provide evidence that, because of their high persistence, the physical dynamics of yields are poorly estimated in samples of the typical length, but at the same time are crucial for model implications. For example, if we were to estimate a VAR using actual inflation instead of decomposing it into trend inflation $\bar{\pi}_t$ and detrended inflation, the VAR would generate long-horizon inflation expectations, and thereby long-horizon expectations of short-term nominal interest rates, that are not nearly volatile enough over our sample (Kozicki and Tinsley, 2001). Using long-horizon survey expectations helps to disentangle transitory dynamics in inflation and real short-term rates from the secular movements as in (5)-(6), most notably the decline in long-horizon inflation expectations during the 1980s and 1990s. These two decompositions are motivated by the substantial fluctuations, shown in Figure 3.2, in expectations 7 to 11 years ahead of nominal 3-month Treasury bill yields, CPI inflation, and the expectations for real short-term interest rates implied by their difference.

In addition to the long-horizon expectations for inflation and the 3-month yield, we use expectations at the 6- and 12-month horizons of the 3-month Treasury bill yield from the

Blue Chip Financial Forecasts in the estimation, and impose that the VAR-implied expectations at the respective horizons are equal to these survey measures plus i.i.d. measurement errors. This assumption implies linear relationships of the form

$$E_t y^{svy}_{1,t+k} = \zeta'_{y,k} X_t + \epsilon^{y,k}_t \tag{12}$$

where $E_t y^{svy}_{1,t+k}$ denotes survey expectations of the 3-month T bill yield k periods ahead, and the coefficients $\zeta_{y,k}$ are functions of the VAR parameters.

An additional refinement of the decomposition (5) is motivated by the fact that CPI inflation, whether measured at monthly or quarterly frequency, is a very noisy process. Duffee (2011), Kim (2009), and Gürkaynak and Wright (2012) have argued that not all fluctuations in inflation are priced in bond yields. We therefore consider realized inflation π_{t+1} as composed of

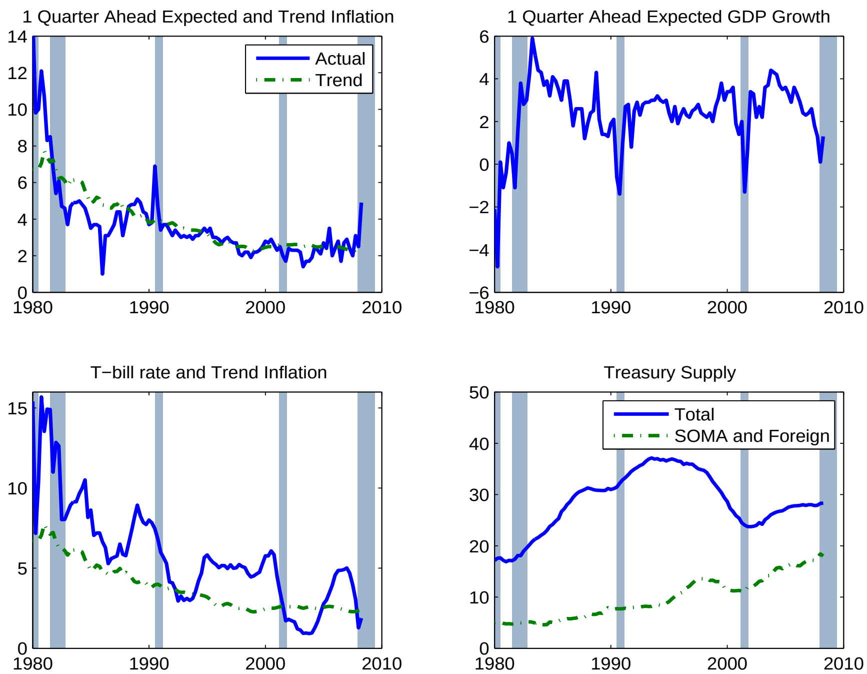
$$\pi_{t+1} = \bar{\pi}_t + \tilde{\pi}_t + \epsilon^{\pi}_{t+1}$$

and that only $E_t \pi_{t+1} = \bar{\pi}_t + \tilde{\pi}_t$ affects bond pricing at date t, whereas the measurement error ϵ^{π}_t does not. In particular, we use one-quarter-ahead survey forecasts of CPI inflation, denoted $E^{svy}_t \pi_{t+1}$ and assume that $E^{svy}_t \pi_{t+1} = \tilde{\pi}_t + \bar{\pi}_t$, thereby identifying $\tilde{\pi}_t$ conditional on an estimate of $\bar{\pi}_t$. Similarly, we use one-quarter ahead survey forecasts of real GDP growth as measure for real activity.

As will be discussed below, we seek to identify shocks that resemble the quantitative easing programs undertaken in recent years by several central banks. Usually, the main source of variation of the amount and duration of Treasury debt held by private investors is the Treasury itself. However, variations associated with fiscal policy actions affect macroeconomic variables importantly through changes in government spending and taxation. By contrast, variations in Treasury debt held by private investors engineered by central banks are not associated with fiscal measures and can therefore be expected to have different macroeconomic effects than variations due to fiscal policy. To allow us to disentangle these different sources of variations in Treasury debt held by private investors, we include two measures. The first, t_t, is the total amount of marketable debt outstanding, whereas the second, s_t, is the amount of Treasuries held by the Federal Reserve's System Open Market Account (SOMA) and foreign official institutions.[7] Both are expressed as percent of nominal GDP. The time series of the elements of x_t are shown in Figure 3.2. A fuller description

[7] The series of foreign official holdings is described in Beltran et al. (2012).

Figure 2: Stationary variables in the state space



of our state space model is provided in Appendix B, and further details on the data we use in Appendix C.

3.3 Identifying Treasury supply shocks

[Very rough and incomplete]

As discussed before, we would like to separately identify exogenous innovations to Treasury supply originating from fiscal policy on the one hand, and exogenous innovations to Federal Reserve and foreign holdings of Treasury securities on the other. Blanchard and Perotti (2002) proposed an identification strategy that takes account of endogenous responses of taxes and spending to output. Their approach to identification has been subject to a number of criticisms, one important of which is the fact that agents in the economy may be aware of the exogenous fiscal shocks before the econometrician observes them. However, recent work by Leeper et al. (2012) and by Caldara and Kamps (2013) suggests that, in VARs that include more variables than Blanchard and Perotti's 3-equation model, adding measures of fiscal news doesn't qualitatively alter the results concerning dynamic responses to tax and spending shocks. Based on these results, we are adapting Blanchard and Perotti's identification strategy to our setting, in which we include not taxes and spending, but instead the total amount of marketable Treasury debt outstanding. In particular, we use the approximate relationship that

$$t_t = t_{t-1}/N_t + g_t - \tau_t$$

where g_t denotes the ratio of federal government expenditures to GDP, τ_t the ratio of federal tax revenues to GDP, and N_t the gross growth rate of nominal GDP between periods $t-1$ and t. This relationship allows us to convert identifying assumptions for tax and spending shocks separately into identifying assumptions for exogenous Treasury supply shocks induced by fiscal policy. Further details are presented in Appendix D.

The key challenge for identifying exogenous fiscal shocks is that there is clear evidence of contemporaneous causality running in both directions: Real revenues and spending are contemporaneously affected by changes in output and inflation because of the automatic stabilizers and lack of indexation of government wages, and output is contemporaneously affected by government spending and arguably by tax changes. By contrast, we assume that U.S. output and inflation are contemporaneously unaffected by exogenous changes in Federal Reserve and foreign official holdings of Treasury securities, for essentially the same reasons that most of the literature on identifying monetary policy shocks has assumed output and inflation to be contemporaneously unaffected by exogenous interest rate shocks. Hence, in addition to applying the Blanchard-Perotti identification strategy to fiscal shocks, we assume that exogenous innovations to the two monetary policy instruments, Federal Reserve and foreign official holdings and the short-term interest rate, do not contemporaneously affect any of the remaining variables in the VAR.

How to disentangle exogenous innovations to the two monetary policy instruments is a challenging question. For now, we assume a recursive ordering in which Federal Reserve and foreign official holdings of Treasury securities are chosen before the interest rate is determined, but we recognize that this is somewhat arbitrary. In future work we will want to explore alternative identifying assumptions in the spirit of Faust and Rogers (2003).

4 Estimation and Results

The model is estimated over the sample 1980Q1 to 2008Q2. We start the sample only in 1980 because the systematic response of monetary policy to economic conditions is an important element of our factor VAR, and there is strong evidence for a break in this systematic component around 1980. We end the sample just before the intensification of the financial crisis in September 2008 because shortly thereafter the nominal short rate reached the zero lower bound (ZLB), thereby introducing a nonlinearity in short-rate dynamics that our affine term structure model does not capture.[8] However, in discussing our results below, we will discuss the likely implications of our model for the effects of QE at the zero lower bound.

As robustness checks we also investigate three alternative sample periods, including a longer pre-crisis sample of 1971Q4 to 2008Q2 as well as two samples that include the most recent ZLB periods: a shorter ZLB sample from 1980Q1 to 2011Q2, and a longer ZLB sample of 1971Q4 to 2011Q2. The ending date of 2011Q2 is determined by the availability of foreign official holdings of Treasury securities. For the latter two sample periods, we impose the ZLB restrictions using the approximation method proposed by Priebsch (2013), which approximates arbitrage-free yields in Gaussian shadow-rate term structure models based on a second-order cumulant-generating-function expansion.

4.1 Estimation

We conduct a two-step estimation of the model. In the first step the VAR parameters are estimated by OLS, treating linearly interpolated long-horizon survey expectations as perfect measures of the inflation and real rate asymptotes. Information criterion-based tests reveal that, once the time-varying long-run trends are removed, the stationary factors in the system call for a first-order VAR, we therefore set $q = 1$ in our estimation. Subsequently, we hold the VAR parameters obtained in the first step fixed and estimate of the $\mathbb{Q}$-parameters by fitting observed yields, their survey forecasts, and other variables in the measurement equation. In the second step we allow all variables to be measured with errors and back out the latent state variables using the standard Kalman Filter if the ZLB is not imposed and the square root unscented Kalman Filter if it is.

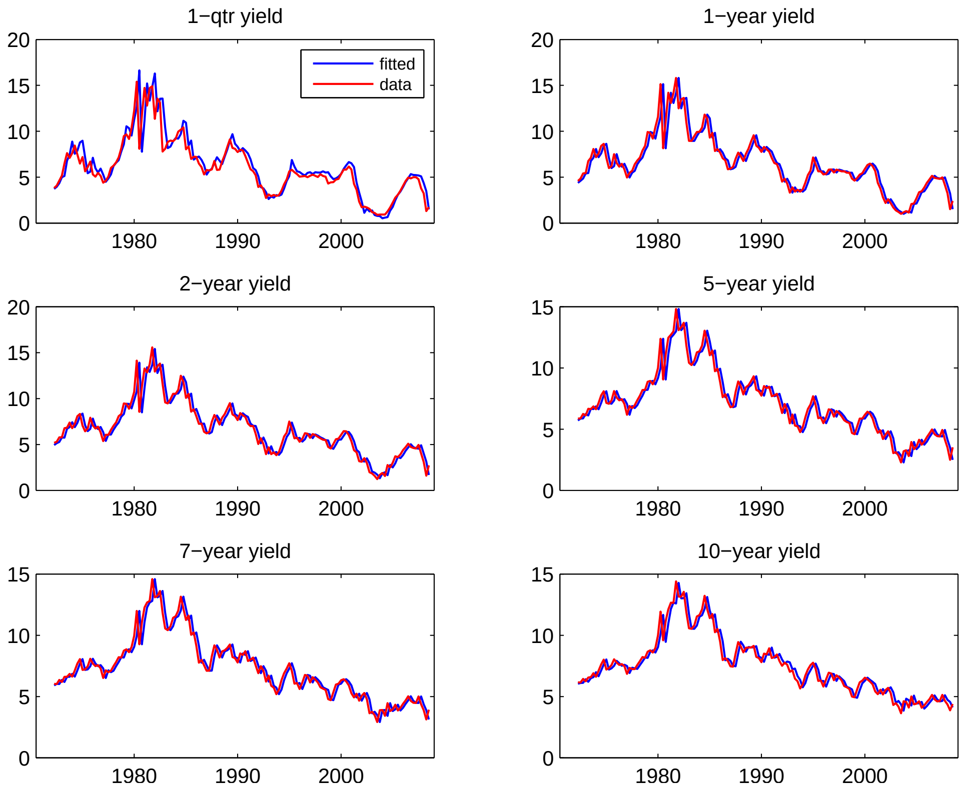
To estimate the model, we use zero-coupon Treasury yields with maturities 1, 2, 3, 7, 10, and 15 years implied by a fitted Svensson (1995) yield curve as described in Gürkaynak, Sack, and Wright (2007).

Figure 3 plots the actual and the model-implied yields at all six maturities. The model

[8] Li and Wei (2013) also end their sample in 2007, just before the onset of the financial crisis.

matches yields very well, although the fit deteriorates a bit at the shortest maturity. This is not surprising as we know that empirically three yield factors are sufficient to explain the bulk of time variations in yields.

Figure 3: Actual and Model-Implied Yields



4.2 Treasury supply shocks and the comovement of yields and macro variables

We now turn to some of the impulse responses to exogenous innovations to SOMA and foreign holdings, the short rate, and to total Treasury supply stemming from fiscal policy. We first focus on the responses of state variables to these shocks, and then on the responses of longer-term yields and term premia.

Figure 4 presents impulse responses to an exogenous increase in foreign official and SOMA Treasury holdings in the amount of 1 percent of GDP (roughly $150 billion at current levels). The upper left panel shows the response of the level of real GDP (the cumulative response of GDP growth). The level of real GDP shows little response to the shock over the first two years after the shock but starts to decline thereafter reaching about 20 basis points below the original level. As shown to the right, inflation rose gradually over time to about 15 basis points above its pre-shock level. As shown in the lower left, the shock leads to a very persistent rise in SOMA and foreign Treasury holdings.In terms of the relationship between these holdings and the traditional short-rate tool of monetary policy, the lower right shows that the short rate declines by about 20 basis points upon impact, and then only gradually rises in response to the increase in inflation.

Figure 4: Impulse response functions to F&S shock

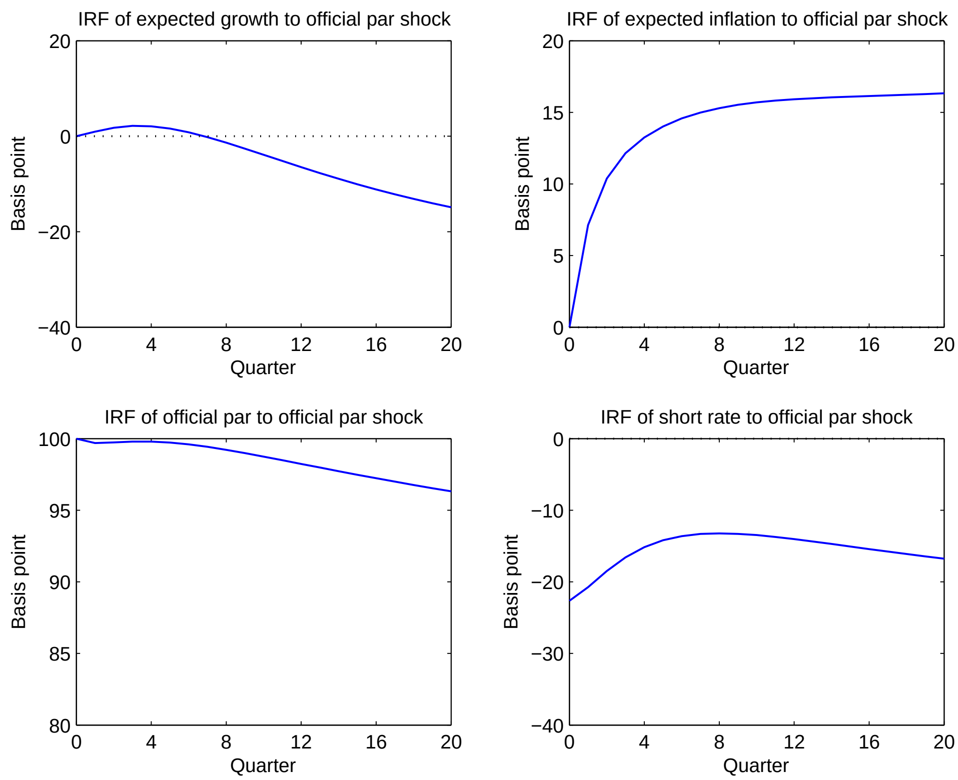
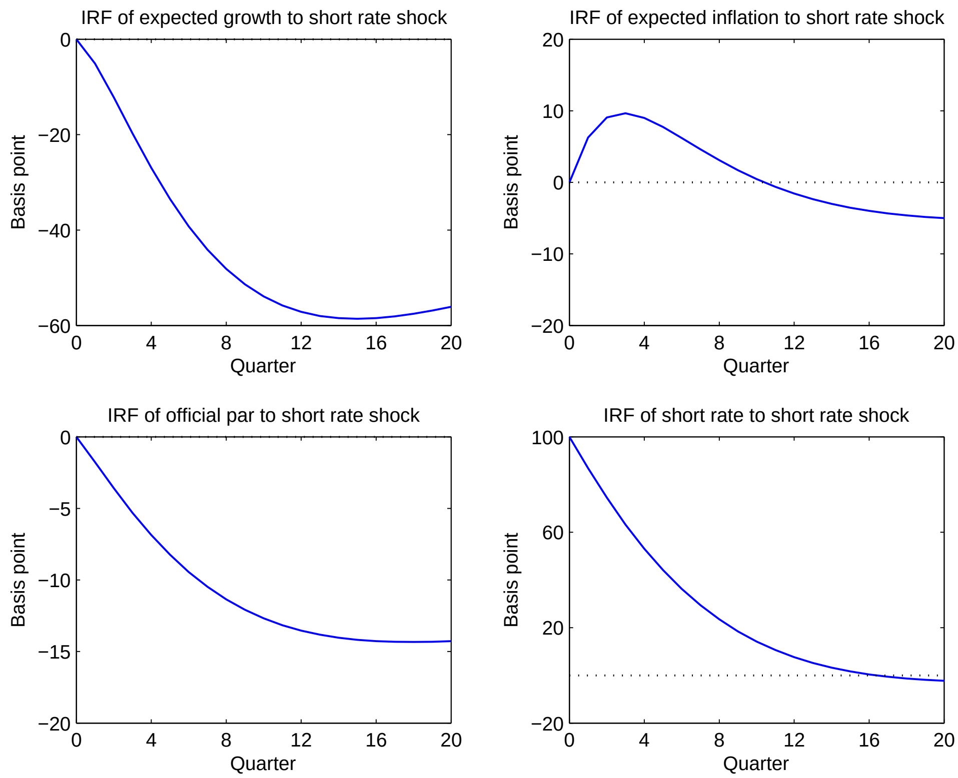


Figure 5 presents impulse responses to a monetary policy innovation of the traditional short-rate variety. In response to a 100 basis point increase (at an annual rate) of the 3-month yield that dies out only gradually, the level of output declines by about 60 basis points over the 10 quarters following the shock before rising. The inflation responses displays a prize puzzle, rising up to 10 basis points 4 quarters after the shock. Finally, the lower left panel indicates that SOMA and foreign holdings act as complements to short-term interest rates by declining by about 15 basis points (about $20 billion at current levels) during the first eight quarters.

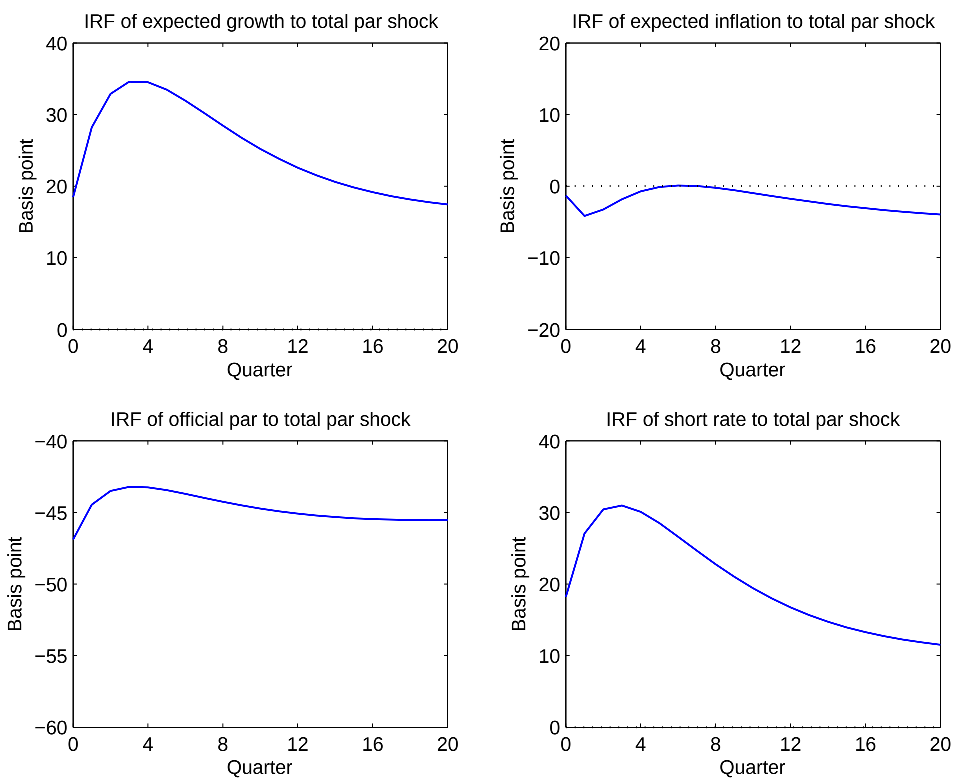
Figure 5: Impulse response functions to short-rate shock



The responses to a fiscal shock are presented in Figure 6. The increase in Treasury supply by 1 percent of GDP could reflect either an increase in spending or a decrease in taxes; since we only include total Treasury supply, we cannot distinguish between these two sources. Output rises upon impact by about 20 basis points and reaches a peak of 35 basis points within four quarters. Inflation show little responses. The increase in output leads to an initial rise in the 3-month T bill yield of about 20 basis points that peaks about three quarter later and dies out gradually thereafter.

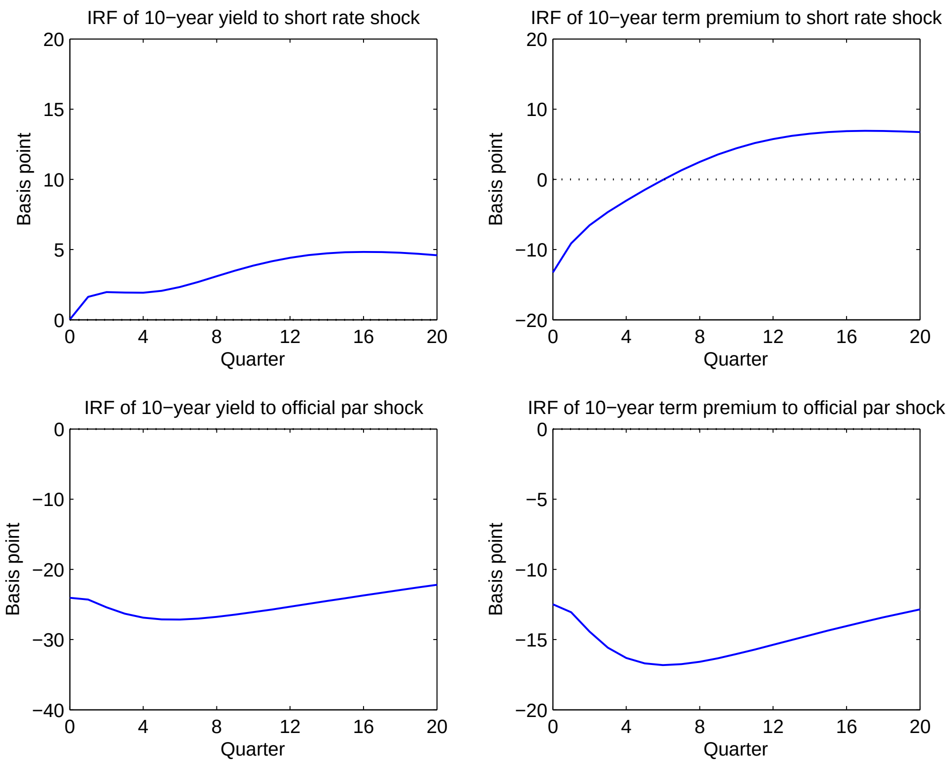
Finally, in Figure 7 we show impulse responses of the 10-year Treasury yield and the associated term premium to the two shocks associated with monetary policy. As the top two panels show, the 10-year yield does not react immediately in response to an exogenous shock to the 3-month yield, as a higher "expectations hypothesis" component of the yield offsets a lower term premium. The 10-year yield rise subsequently by up to 5 basis points,

Figure 6: Impulse response functions to fiscal shock



reflecting almost entirely a rise in the term premium. By contrast, an increase in SOMA and foreign holdings leads to a decline in the 10-year yield that peaks at about 25 basis after 4 quarters, reflecting both lower expected future short rates and a persistent decline in the term premium by about 15 basis points. This estimate is qualitatively similar, but somewhat larger than the estimates reported in Li and Wei (2013).

Figure 7: Impulse responses of 10-year yield and term premium



4.3 Forward Guidance and Term Premium Shocks

The second and the third yield curve factors are hard to interpret economically. Given the affine setup of the model, we could rotate them into factors that have more economic meanings. For example, Gürkaynak, Sack, and Swanson (2005) emphasized that monetary policy affects asset prices and the macroeconomy not only by changing the current stance of policy but also by influencing market expectations of the future path of policy. The "forward guidance" of future monetary policy has become one of the main tools that the Federal Reserve relied on heavily during the most recent financial crisis, as the traditional policy tool, the nominal short rate, became constrained by the zero lower bound.

The other prominent unconventional monetary policy tool used repeatedly during the crisis is asset purchases by the Federal Reserve that are designed to place downward pressures on longer-term Treasury yields, at least partially by reducing the term premium. Nonetheless, as pointed out by Rudebusch, Sack, and Swanson (2007), the existing theoretical and empirical literature provides inconclusive and frequently conflicting answers to the question whether a negative shock to the term premium is expansionary or contractionary. More recently, Kiley (2012) finds that a reduction in term premiums has a stimulative effect on real economic activities but the magnitude of the effect is much smaller than that of a decline in the expected future short-term interest rates.

To shed light on the implications of those two types of shocks, we rotate the state variables such that the last two factors in the VAR now represent the average expected short rate over the next four quarters, $y^1_{t,EH}$, and the 10-year term premium, $y^{10}_{t,TP}$, respectively.

$$z_t = [\bar{\pi}_t,\ \bar{r}_t,\ \tilde{\pi}_t,\ q_t,\ t_t,\ s_t,\ \tilde{\mathcal{P}}^1_t,\ y^1_{t,EH},\ y^{10}_{t,TP}]' \tag{13}$$

We then calculate impulse responses of the macroeconomy and yields at different maturities to shocks to those two shocks, plotted in Figures 8 and 9.

Figure 8 shows that a 100 basis point positive exogenous increase in the average expected future short rates over the next 4 quarters leads expected inflation to decline over the next 10 quarters by up to 20 basis points, while the level of output shows a counterintuitive sharp rise shortly after the shock.

Figure 9 shows that shocks to the 10-year term premium dissipate fairly quickly and largely disappears after 4 quarters. Nonetheless, it appears to lead to small increases in both inflation and the level of output. This is despite a notably upward shift in the entire yield curve, suggesting that these shocks might be at least partially proxying for other fundamental shocks that are favorable to the economy.

4.4 Out-of-Sample Analysis

We estimate the model using data up to the eve of the recent financial crisis. However, this model can also be used to understand the development after the onset of the crisis. To do this, we hold fixed the parameter estimates and use the Kalman filter to infer the values of the latent state variables and the shocks to those variables from macroeconomic variables, yields, and survey forecasts observed during and after the crisis.

Figure 8: Impulse responses to 1-year expected short rate shocks

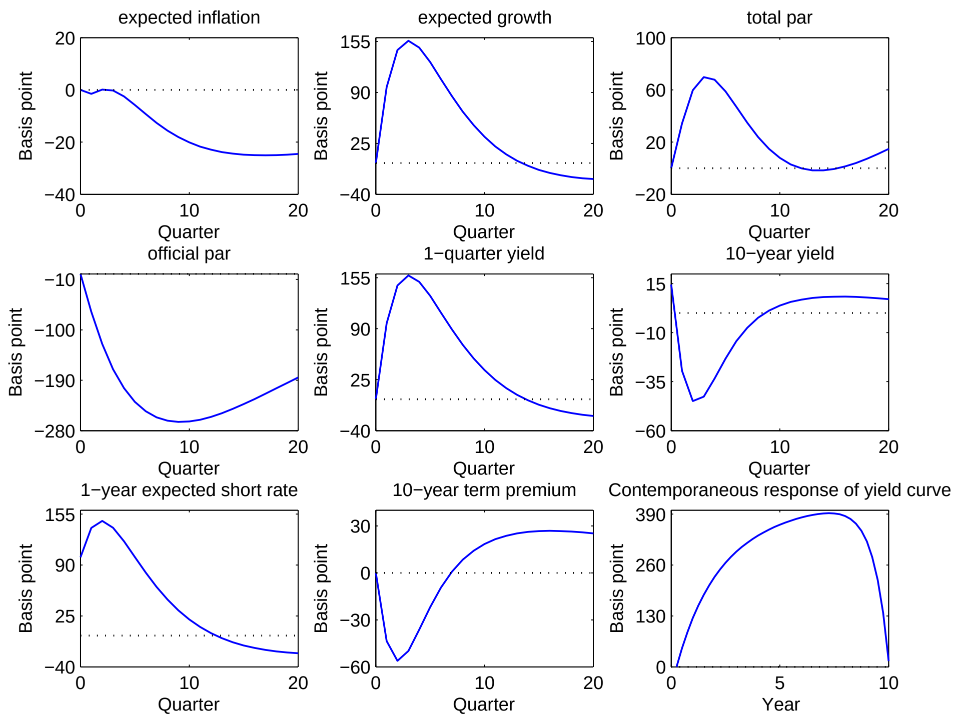
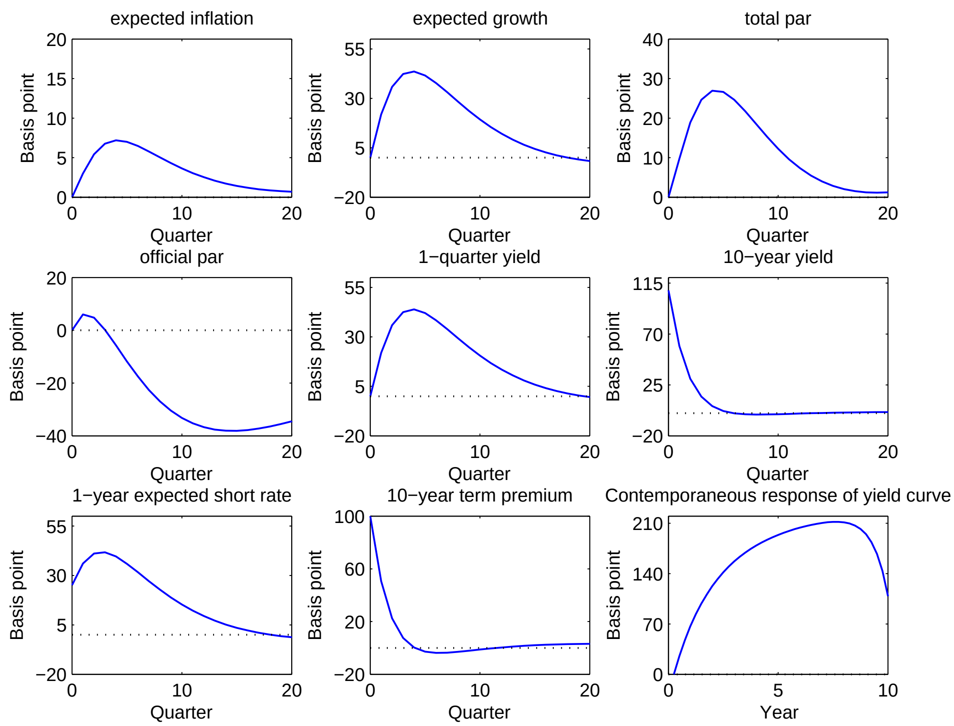
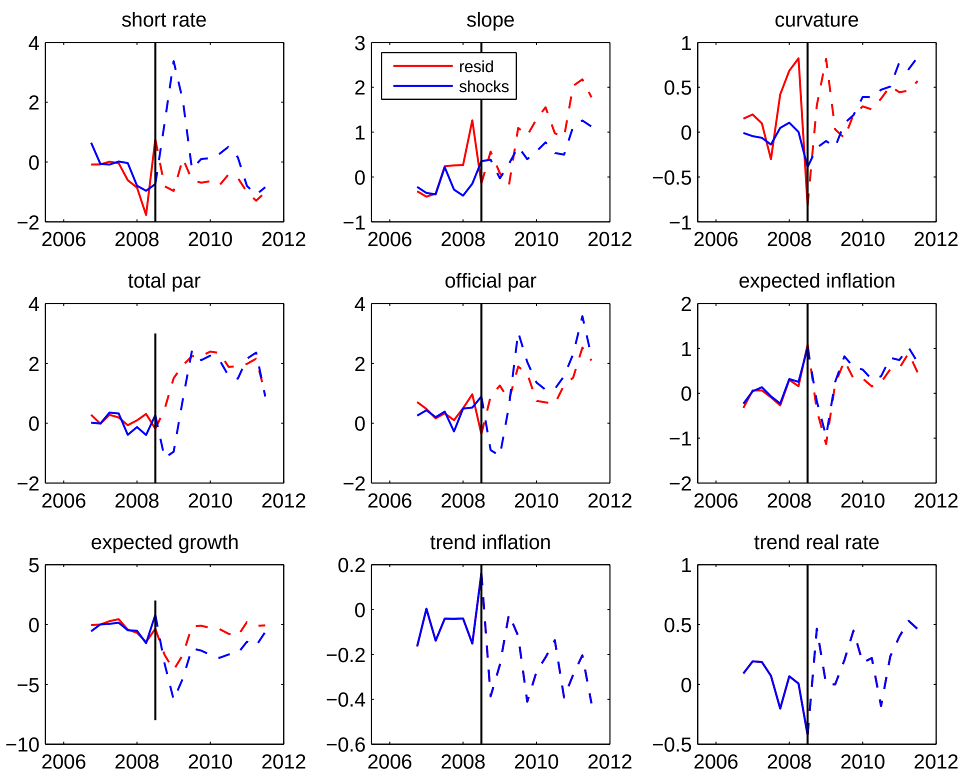


Figure 9: Impulse responses to 10-year term premium shocks



The estimated reduced-form residuals and structural shocks are plotted in Figure 10. The solid lines to the left of the vertical lines denoting 2008Q2 represent in-sample estimates, whereas the dashed lines to the right are out-of-sample estimates. The model interprets the crisis period as accompanied by large positive shocks to the nominal short rate, reflecting the heavy constraints on monetary policy by the zero lower bound, and repeated negative shocks to trend inflation, while shocks to the trend real rate are more symmetrically distributed around zero. A large positive shock to the foreign official and SOMA holdings of Treasury securities, mostly the latter, more than offsets the effect of a large increase in total Treasury debt outstanding caused by the recession. Near-term growth expectations also experienced mostly negative shocks, while near-term inflation expectations are more stable.

Figure 10: Out-of-Sample Analysis: Residuals and Structural Shocks



5 Conclusions

Still to be done:

- Are identified F/S, fiscal shocks in accordance with narrative record? (Favero and Giavazzi, 2012)
- Identify exogenous changes to Treasury maturity composition from historical records. Combine SVAR and narrative approaches to identification in the manner of Stock and Watson (2012), Mertens and Ravn (2013).
- Revisit assumption that SOMA and foreign holdings don't respond contemporaneously to exogenous short-rate shocks.
- Simulate yields over period since 08Q2, decompose into contributions from SOMA purchases, forward guidance, fiscal.
- Bootstrapping the standard errors

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A Specification of the affine term structure model

Rewrite the PC factors in terms of the detrened PC1, the other two PC factors, and the trend variables as $\tilde{\mathcal{P}} = \left[\tilde{\mathcal{P}}_t^1\ \mathcal{P}_t^2\ \mathcal{P}_t^3\ \bar{\pi}_t\ \bar{r}_t\right]'$. It can be seen that Equation (1) is consistent with the following $\mathbb{Q}$-VAR(1) on $\tilde{\mathcal{P}}$:

$$\tilde{\mathcal{P}}_t = \mu^{\mathbb{Q}}_{\tilde{\mathcal{P}}} + \Phi^{\mathbb{Q}}_{\tilde{\mathcal{P}}}\tilde{\mathcal{P}}_{t-1} + \Sigma^{\mathbb{P}}_{\tilde{\mathcal{P}}}\epsilon^{\mathbb{Q}}_{\tilde{\mathcal{P}},t}$$

where

$$\mu^{\mathbb{Q}}_{\tilde{\mathcal{P}}} = \begin{bmatrix} \mu^{\mathbb{Q}}_{\mathcal{P}} \\ 0 \\ 0 \end{bmatrix}, \; \Phi^{\mathbb{Q}}_{\tilde{\mathcal{P}}} = \begin{bmatrix} \phi^{\mathbb{Q}}_{\mathcal{P}} & 0_{3\times 1} & 0_{3\times 1} \\ 0_{1\times 3} & \phi^{\mathbb{Q}}_{\mathcal{P},11} & 0 \\ 0_{1\times 3} & 0 & \phi^{\mathbb{Q}}_{\mathcal{P},11} \end{bmatrix}$$

The nominal short rate loads only on $\tilde{\mathcal{P}}$; as a result only those factors are priced in the nominal Treasury market:

$$\log M_{t+1} = -y^1_t - \frac{1}{2}\Lambda_t'\Lambda_t - \Lambda_t'\epsilon^{\mathbb{Q}}_{\tilde{\mathcal{P}},t+1}, \tag{14}$$

with the price of risk parameters determined by the parameters governing the $\mathbb{P}$- and the $\mathbb{Q}$-VARs on $\tilde{\mathcal{P}}$:

$$\lambda_t = \lambda_0 + \lambda_1 \tilde{\mathcal{P}}_t \tag{15}$$

where

$$\begin{aligned} \lambda_0 &= \left(\Sigma^{\mathbb{P}}_{\tilde{\mathcal{P}}}\right)^{-1}\left(\mu^{\mathbb{P}}_{\tilde{\mathcal{P}}} - \mu^{\mathbb{Q}}_{\tilde{\mathcal{P}}}\right) \\ \lambda_1 &= \left(\Sigma^{\mathbb{P}}_{\tilde{\mathcal{P}}}\right)^{-1}\left(\Phi^{\mathbb{P}}_{\tilde{\mathcal{P}}} - \Phi^{\mathbb{Q}}_{\tilde{\mathcal{P}}}\right) \end{aligned}$$

B The state space model with survey information

The model is specified at the quarterly frequency. In the estimation of the model, we use 3-month and 1-, 2-, 3-, 7-, 10-year Treasury yields. The vector of observables Y_t therefore consists of

$$\begin{aligned} Y_t = [\; & y^1_t,\ y^4_t,\ y^8_t,\ y^{12}_t,\ y^{28}_t,\ y^{40}_t,\ E^{BC}_t[\pi_{t+7\to 11}],\ E^{BC}_t[r_{t+7\to 11}], \\ & E^{BC}_t[\pi_{t+1}],\ E^{BC}_t[q_{t+1}],\ t_t,\ s_t,\ E^{BC}_t[y^1_{t+2}],\ E^{BC}_t[y^1_{t+4}]\ \;]', \end{aligned} \tag{16}$$

where $E^{BC}_t[\pi_{t+7\to 11}]$ and $E^{BC}_t[y^1_{t+7\to 11}]$ denote the Blue Chip forecasts of CPI inflation and the 3-month T bill yield at the longest horizon (the projected average over the horizon roughly 7 to 11 years ahead), and $E^{BC}_t[y^1_{t+2}]$ and $E^{BC}_t[y^1_{t+4}]$ are the Blue Chip forecasts of the 3-month T bill yield 2 and 4 quarters ahead. We treat the long-run survey expectations $E^{BC}_t[\pi_{t+7\to 11}]$ and $E^{BC}_t[y^1_{t+7\to 11}]$ as if they had a constant forecast horizon 25 to 44 quarters ahead, and calculate the model-implied expectation of average inflation over this horizon as $\zeta_\pi' X_t$ with

$$\zeta_\pi = 0.05(\iota_4 + \iota_9)\Phi^{25}(\sum_{i=0}^{19}\Phi^i)$$

where ι_j selects the j-th element of X_t.

With these assumptions, the state space model consists of the transition equation given by (9) and a measurement equation

$$Y_t = A + BX_t + e_t$$

given by

$$\begin{bmatrix} y_t^1 \\ \vdots \\ y_t^{40} \\ E_t^{BC}[\pi_{t+7\to 11}] \\ E_t^{BC}[y^1_{t+7\to 11}] \\ E_t^{BC}[\pi_{t+1}] \\ E_t^{BC}[q_{t+1}] \\ t_t \\ s_t \\ E_t^{BC}[y^1_{t+2}] \\ E_t^{BC}[y^1_{t+4}] \end{bmatrix} = \begin{bmatrix} \\ A_y \\ \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} & & & & & & & & \\ & & & & B_y & & & & \\ & & & & & & & & \\ & & & & \zeta'_{\bar{\pi}} & & & & \\ & & & & \zeta'_{\bar{r}} & & & & \\ 1 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ & & & & \zeta_y^{2'} & & & & \\ & & & & \zeta_y^{4'} & & & & \end{bmatrix} \begin{bmatrix} \bar{\pi}_t \\ \bar{r}_t \\ \tilde{\pi}_t \\ q_t \\ t_t \\ s_t \\ \tilde{y}_t^1 \\ y_t^2 \\ y_t^3 \end{bmatrix} + \begin{bmatrix} \epsilon_t^1 \\ \vdots \\ \epsilon_t^{40} \\ \epsilon_t^{\bar{\pi}} \\ \epsilon_t^{\bar{r}} \\ \epsilon_t^{\pi,1} \\ \epsilon_t^{q,1} \\ \epsilon_t^{t} \\ \epsilon_t^{s} \\ \epsilon_t^{y,2} \\ \epsilon_t^{y,4} \end{bmatrix} \quad (17)$$

where A_y and B_y are the vector and matrix of the stacked coefficients a_n and b_n from Equation (3), The corresponding measurement errors are denoted by ϵ.

C The data

[To be completed]

The series of inflation expectations consists of three different pieces. Until 1981:Q1, the series is an estimated step function based on the changepoint model developed in Kozicki and Tinsley (2001). From 1981:Q2 until 1988:Q4, the series is based on the Hoey survey of bond market participants, which was conducted on a quarterly basis by Richard Hoey, an economist at Drexel Burnham Lambert. Participants in this survey were polled for their expectation of CPI inflation over the second five years of a 10-year horizon. From 1989:Q4 onwards the series is based on the expectations for the average CPI inflation rate roughly 7 to 11 years ahead from the Blue Chip Financial Forecasts. Because the Financial Forecasts poll respondents only twice a year for their long-horizon forecasts, we interpolate the data to quarterly frequency. For the years from 1986 on, we use the long-horizon forecasts of the 3-month T bill yield from the Financial Forecasts, and likewise interpolate to quarterly frequency. Before 1986 no such long-horizon forecasts are available, and we treat them as missing observations in the estimation.

D Identification and estimation of the VAR

Following the notation used in (9), let v_t denote the vector of reduced-form residuals, and ε_t the vector of structural innovations of which we seek to identify several elements. As stated in the main text, we assume that the asymptote of inflation $\bar{\pi}_t$ and of the one-period real rate $\bar{r}_t$ follow univariate random walks with innovations $\varepsilon_t^{\bar{\pi}}$ and $\varepsilon_t^{\bar{r}}$.

We are interested in identifying the structural shocks to total Treasury supply, including privately-held Treasuries, and those to Treasury holdings by the SOMA and foreign official institutions. Following Blanchard and Perotti (2002), we assume that fiscal policy cannot respond contemporaneously to macroeconomic developments except by the automatic stabilizers embedded in the tax and spending policies in place. Hence, the reduced-form innovations to total Treasury supply are composed of a response to current shocks to economic activity q and inflation π as implied by the automatic stabilizers, and any exogenous fiscal policy shocks that are unrelated to current macroeconomic conditions. Note in particular that total Treasury supply is assumed to be contemporaneously unaffected by monetary policy, be that SOMA asset holdings or the one-period interest rate (where we follow the convention of assuming no contemporaneous response of real activity and inflation to innovations to $y_{1,t}$). With these assumptions, the contemporaneous relationship between the reduced-form innovations v_t^t, v_t^q, and v_t^π to Treasury supply, real activity, and inflation respectively, and the structural fiscal (Treasury supply) shock ε_t^t is

$$v_t^t = \eta^{t,\pi} v_t^\pi + \eta^{t,q} v_t^q + \varepsilon_t^t$$

where the coefficients $\eta^{t,x}$ can be constructed as $\eta^{t,x} = \eta^{\tau,x} - \eta^{g,x}$ from the underlying calibrated parameters in the equations for log real taxes τ and log real spending g

$$\begin{aligned} v_t^\tau &= \eta^{\tau,\pi} v_t^\pi + \eta^{\tau,q} v_t^q + \varepsilon_t^\tau \\ v_t^g &= \eta^{g,\pi} v_t^\pi + \eta^{g,q} v_t^q + \varepsilon_t^g \end{aligned}$$

Constructing the coefficients $\eta^{t,\pi}$ and $\eta^{t,q}$ in the manner of Blanchard and Perotti (2002) is critical because inflation and real activity are assumed to be contemporaneously affected by fiscal policy and hence by ε_t^t.

By contrast, SOMA (and foreign official) Treasury holdings f_t are also assumed to respond contemporaneously to real activity and inflation, whereas they are not assumed to affect real activity and inflation, in analogy to the conventional assumption in the literature that monetary policy shocks (in the form of structural shocks to the short-term interest rate) do not affect these variables contemporaneously. The relationships between the reduced-form residuals v_t and the structural innovations ε_t can thus be written as $v_t = \boldsymbol{\eta}\varepsilon_t$ with

$$\begin{bmatrix} v_t^{\bar{\pi}} \\ v_t^{\bar{r}} \\ v_t^{\pi} \\ v_t^{q} \\ v_t^{t} \\ v_t^{s} \\ v_t^{y} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ ? & ? & 1 & 0 & ? & 0 & 0 \\ ? & ? & ? & 1 & ? & 0 & 0 \\ ? & ? & \eta^{t,\pi} & \eta^{t,q} & 1 & 0 & 0 \\ ? & ? & ? & ? & 0 & 1 & 0 \\ ? & ? & ? & ? & ? & ? & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_t^{\bar{\pi}} \\ \varepsilon_t^{\bar{r}} \\ \varepsilon_t^{\pi} \\ \varepsilon_t^{q} \\ \varepsilon_t^{t} \\ \varepsilon_t^{s} \\ \varepsilon_t^{y} \end{bmatrix} \tag{18}$$

The parameters $\eta^{t,\pi}$ and $\eta^{t,q}$ are calibrated based on the values for these parameters reported in Perotti (2004). The parameters denoted with "?" are estimated by instrumental variables. Specifically, the structural residuals $\varepsilon_t^{\bar{\pi}}$ and $\varepsilon_t^{\bar{r}}$ are simply the first differences of the series $\bar{\pi}_t$ and $\bar{r}_t$; the unknown parameters in the third row of the matrix are estimated by regressing v_t^{π} on $\varepsilon_t^{\bar{\pi}}$, $\varepsilon_t^{\bar{r}}$, and v_t^t,, using ε_t^t as instrument etc.