Maturity Structure and Supply Factors in Japanese Government Bond Markets

Ichiro Fukunaga, Naoya Kato, and Junko Koeda

Using our constructed database on the amount outstanding of Japanese Government Bonds (JGBs) categorized by holder and remaining maturity, we examine the effects of changes in the holders and maturity structures on the term structure of interest rates and the risk premium on long-term bonds. Both approaches using single-equation regressions and a term structure model confirm that the net supply of JGBs, the issuance (supply) by the government minus the demand by the preferred-habitat investors including the Bank of Japan (BOJ), had significant effects on long-term interest rates. The regression approach implies that the net supply effects were stronger in the zero interest rate periods, while this relationship was not found using the model approach. We also calculate the net supply effects of the BOJ's JGB purchases as part of its Quantitative and Qualitative Monetary Easing and compare the results with those obtained from a simple event-study analysis.

Keywords: Japanese Government Bonds; Term structure of interest rates; Preferred habitat; Unconventional monetary policy
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I. Introduction

After the recent global financial crisis, many central banks in advanced economies lost their conventional monetary policy tools near the zero lower bound on short-term interest rates. To provide further monetary accommodation, they have moved to unconventional tools, such as forward guidance and government bond purchases, aimed at affecting longer-term interest rates.[1] The Bank of Japan (BOJ) introduced these unconventional tools long before the global financial crisis,[2] but since April 2013, it has implemented more aggressive monetary easing measures, the "Quantitative and Qualitative Monetary Easing (QQE)," to overcome prolonged deflation and achieve the price stability target.[3] Under the QQE, the pace of the BOJ's purchases of Japanese Government Bonds (JGBs) has accelerated and the remaining maturity of its JGB purchases has been extended significantly, while clear forward guidance on the continuation of the policy has been provided.

In general, a central bank's purchases of government bonds are considered to lower long-term interest rates through three channels: the signaling channel, the scarcity channel, and the duration channel.[4] The signaling channel captures the role of bond purchases as a credible commitment by the central bank to keep interest rates low and thus serves as a complement to forward guidance. The bond purchases may support the credibility because market participants may perceive that it would be difficult for the central bank to raise rates rapidly with a large amount of long-term assets on its balance sheet. The scarcity channel captures the effect of a shortage of bonds with particular maturities available for trading, caused by the central bank's purchase of these particular bonds, on the corresponding yields. The duration channel captures a downward shift of the entire yield curve through a decrease in the price of duration risk (risk premium) caused by the central bank's purchases of longer-term bonds. Among these channels, the signaling channel had been considered to be the main channel until recently. Theoretically, according to the expectations hypothesis, changes in the supply of bonds should affect yields only to the extent that expectations of future short-term interest rates are changed, that is, only through the signaling channel.[5] Similarly,

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1. Many central banks in advanced economies purchased private as well as public assets, especially in the earlier part of the post-crisis period, mainly to restore the functioning of financial markets and intermediation rather than provide monetary accommodation.
2. The BOJ first introduced a zero interest rate policy in 1999 after a domestic banking crisis, and then implemented a quantitative monetary easing policy from 2001 through 2006 with forward guidance on the continuation of the policy. After the global financial crisis in 2008, the BOJ again cut the policy rate to virtually zero and established an asset purchase program under its "Comprehensive Monetary Easing."
3. The BOJ's price stability target was set at 2 percent in terms of the year-on-year rate of change in the consumer price index in January 2013. Before that, the "price stability goal in the medium to long term" was set at 1 percent (in a positive range of 2 percent or lower). A crucial feature of the QQE is that it aims to raise inflation expectations through a strong and clear commitment to achieving the price stability target as well as massive purchases of JGBs, and thereby to lower *real* interest rates. In this paper, however, our analyses focus on nominal rather than real interest rates. For an analysis of real interest rates and the inflation risk premia in Japan, see for instance Imakubo and Nakajima (2015).
4. This classification follows IMF (2013). Krishnamurthy and Vissing-Jørgensen (2011) classify the channels through which the Federal Reserve's quantitative easing policies affected a broader range of long-term interest rates, including the interest rates on corporate bonds and real interest rates. Joyce *et al*. (2012) survey the theoretical and empirical literature on the transmission mechanisms of quantitative easing policies.
5. Gürkaynak and Wright (2012) discuss various implications and limitations of the expectations hypothesis and

Eggertsson and Woodford (2003) argue that the central bank's open-market operations would be irrelevant if they did not change the expected future conduct of monetary or fiscal policy. Empirically, at least for Japan, existing studies on the earlier quantitative easing policy imply that the effects through the scarcity and duration channels, often combined and referred to as the portfolio rebalance channel, were smaller or less clear-cut than those through the signaling channel.[6]

Recent studies motivated by the implementation of large-scale bond purchases by many central banks in advanced economies, however, pay more attention to the scarcity and duration channels. A theoretical background of the scarcity channel is the preferred-habitat theory, first set forth by Modigliani and Sutch (1966), which assumes that bond markets are segmented due to the existence of "preferred-habitat" investors with preferences for particular maturities and thus the interest rate for a given maturity is determined by the supply and demand of bonds with that particular maturity. Based on this assumption, Vayanos and Vila (2009) construct a model of the term structure of interest rates in which the term structure and bond risk premia are determined through the interaction between the preferred-habitat investors and risk-averse arbitragers and thereby the duration channel generated by the arbitragers' aversion to duration risk as well as the scarcity channel is incorporated.[7] Two important predictions of this theoretical framework are as follows: (i) the "net supply" of more and longer-term government bonds, which is defined as the issuance (supply) by the government minus the demand by the preferred-habitat investors, leads to a higher risk premium and steeper yield curve; and (ii) the above "net supply" effects are stronger when the arbitragers' risk aversion is higher. In the real world, the preferred-habitat investors include long-term investors such as pension funds and insurance companies as well as central banks that implement large-scale bond purchases as an unconventional monetary policy tool. The preferred-habitat theory can explain a variety of episodes in which a shock to the net supply of government bonds affected long-term interest rates,[8] including the recent bond purchases by central banks in advanced economies, and appears to be increasingly accepted by policymakers as well as academic economists.[9]

In this paper, we examine how the net supply of JGBs affected the term structure of interest rates and the risk premium on long-term JGBs using Japanese monthly data from January 1992 to September 2014, which includes the period after the QQE was introduced.[10] First we construct a database on the amount outstanding of JGBs categorized by holder and remaining maturity, from which various net supply measures are

standard affine term structure models.

6. For details, see for instance Ugai (2007) and Ueda (2012a, c).
7. The model of Vayanos and Vila (2009) can be positioned within the literature on "limits of arbitrage" surveyed by Gromb and Vayanos (2010) that investigates how costs faced by arbitragers, that is, shocks to the demand and supply of government bonds with particular maturities absorbed by the risk-averse arbitragers in this model, can prevent them from eliminating mispricings.
8. Greenwood and Vayanos (2010) present two recent episodes that strongly support the relevance of the preferred-habitat theory: the U.K. pension reform of 2004 and the U.S. Treasury buyback program of 2000–01.
9. Kohn (2009) and Yellen (2011) mentioned the preferred-habitat theory as a key mechanism behind the effects of the Federal Reserve's large-scale asset purchases on long-term interest rates.
10. Our sample period does not include the period after the expansion of the QQE was announced on October 31, 2014.

calculated. To examine the net supply effects, we use two approaches: single-equation regressions and a term structure model. The former approach can flexibly control various factors that affect long-term interest rates, while the latter is more rigorously related to the preferred-habitat theory. Both approaches confirm that the net supply effects are statistically significant, which is consistent with prediction (i) above on the preferred-habitat theory. Moreover, the regression approach implies that the net supply effects were stronger when the arbitragers' risk aversion was higher, and this relationship is incorporated in the model approach, consistently with prediction (ii) above.[11] However, while the regression approach implies that the net supply effects were stronger in the zero interest rate periods when the arbitragers' risk aversion was relatively high, this relationship was not found using the model approach, because the model estimation results imply that the effects were weaker in the zero interest rate periods given the arbitragers' risk aversion. Finally we calculate the net supply effects of the BOJ's JGB purchases as part of the QQE from both approaches and compare the results with those obtained from a simple event-study analysis.

Our regression approach is closely related to Greenwood and Vayanos (2014), who examine the effects of the supply and maturity structure of U.S. government debt on bond yields and returns based on the preferred-habitat theory. Unlike us, however, they focus on the gross supply of government debt and do not use data on the bonds held by preferred-habitat investors to construct net supply measures. Kuttner (2006) examines the effects of changes in the Federal Reserve's holdings of long-term securities on the holding-period excess returns, and Chadha *et al.* (2013) examine the effects of the average maturity of U.S. Treasury bonds held outside the Federal Reserve on forward long-term yields and the term premium. D'Amico *et al.* (2012) examine the effects through the scarcity channel and duration channel on several components of long-term U.S. Treasury yields. Besides these, many regression analyses estimate the effects of government debt or a central bank's purchases of government debt on long-term interest rates, including Laubach (2009) for U.S. government debt[12] and D'Amico and King (2013) and Meaning and Zhu (2011) for the Federal Reserve's purchases,[13] although they examine the effects through broader channels than those motivated by the preferred-habitat theory.[14] Meanwhile, our term structure model approach is based

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11. Strohsal (2013) examines the net supply effects under time-varying risk aversion using German daily data and obtains results consistent with prediction (ii) above on the preferred-habitat theory.
12. Following Laubach (2009), Kameda (2014) estimates the effects of government debt and budget deficits on long-term JGB yields using Japanese data. Ichiue and Shimizu (2012) examine the determinants of long-term bond yields, including fiscal conditions, using panel data for 10 developed countries.
13. To examine the effects of a central bank's purchases of government bonds on long-term interest rates, many studies use event-study methodologies: for instance, Gagnon *et al.* (2011), Krishnamurthy and Vissing-Jørgensen (2011), Swanson (2011), Neely (2010), and Cahill *et al.* (2013). Related studies using Japanese data include Ueda (2012b) and Lam (2011). Some of these studies use regression analyses as well as event-study analyses.
14. In this relation, Krishnamurthy and Vissing-Jørgensen (2012) show that when the supply of the U.S. government debt is low, the value that investors assign to the liquidity and safety attributes offered by Treasury bonds is high, and thus the yield on Treasuries is low relative to the yield on corporate bonds that offer less liquidity and safety. This implies that the preferred-habitat theory can be applied to not only within government bonds market but also between government bonds and other bond markets. Meanwhile, Greenwood, Hanson, and Stein (2010) show that when the supply of long-term bonds by the government is low, private firms issue a larger fraction of their debt long term given the existence of preferred-habitat investors for long-term bonds.

on a model extended from Hamilton and Wu (2012a),[15] who incorporate preferred-habitat investors into an arbitrage-free term structure model as a discrete-time version of Vayanos and Vila (2009) and estimate it using U.S. data. Also motivated by the preferred-habitat theory, Li and Wei (2013) estimate an arbitrage-free term structure model with U.S. Treasury and mortgage-backed securities supply factors as well as observable yield factors and evaluate the effects of the Federal Reserve's large-scale asset purchase programs. While all the papers mentioned above conduct empirical analyses using U.S. data, similar analyses using recent Japanese data including the period under the QQE are very limited.[16] Moreover, similar analyses using data on private preferred-habitat investors as well as a central bank are limited even in the U.S.[17]

The remainder of the paper is organized as follows. First, we discuss our JGB data in Section II. Then, the results from a regression approach and a term structure model approach are reported in Sections III and IV, respectively. Finally, we show the results calculated from both approaches as well as a simple event-study analysis for the net supply effects of the BOJ's JGB purchases in Section V. Section VI concludes. Appendix 1 explains the construction of the JGB database discussed in Section II and Appendix 2 describes the term structure model used in Section IV.

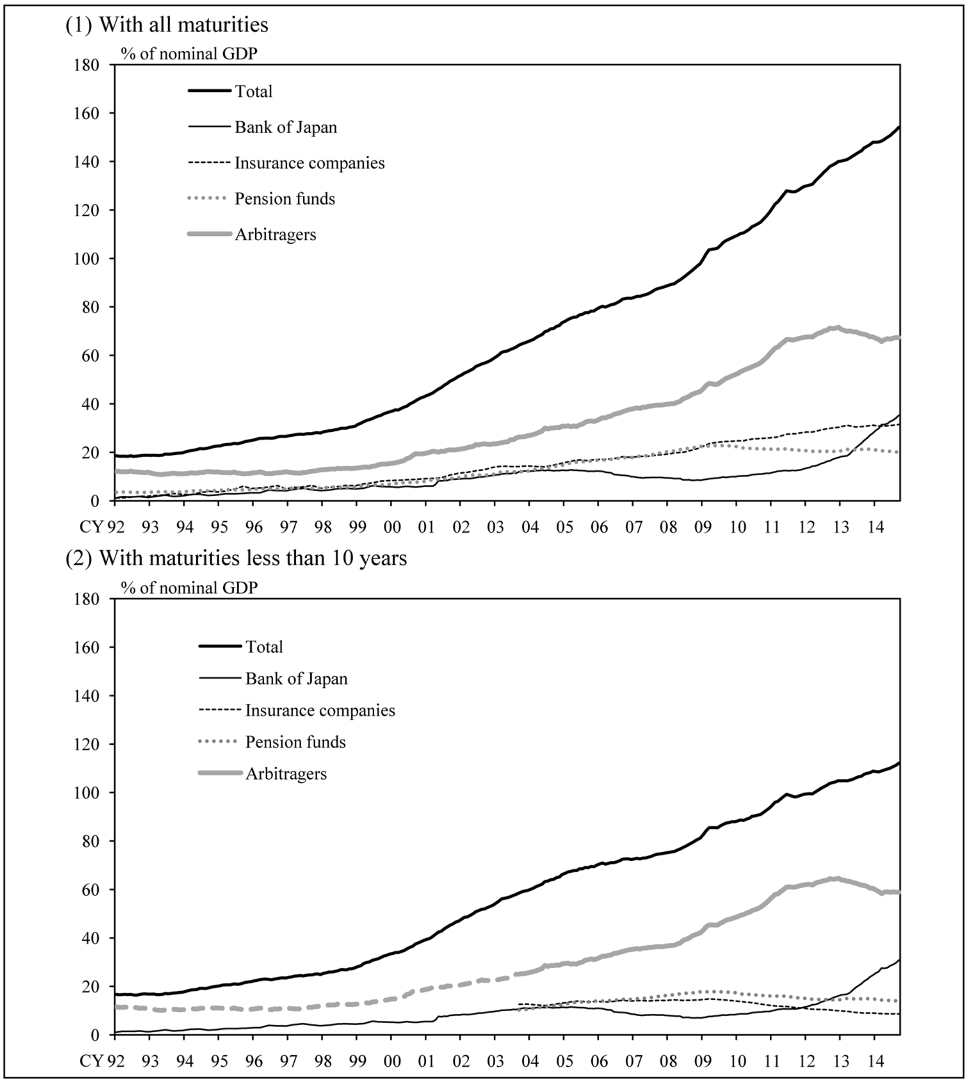
II. Data on JGB Outstanding and Yields

In this section, we discuss data on the holders and maturity structures of JGBs and the term structure of JGB yields. The details of the data construction with regard to the amount outstanding of JGBs categorized by holder and remaining maturity are described in Appendix 1.

Figure 1 presents the amount outstanding of marketable fixed-rate JGBs[18] divided by nominal GDP from January 1992 to September 2014, which corresponds to the main sample period of our analyses in the following sections.[19] The upper panel shows those with all maturities, and the lower panel shows those with remaining maturities less than 10 years. Besides the total amounts issued by the government, Figure 1

15. The model used in this paper is the same as in Koeda (2015), which extends Hamilton and Wu (2012a) by allowing the coefficients in yield-curve factor dynamics as well as the prices of risk to change at the zero lower bound regime, using latent factors instead of observable factors, and so on. In this paper, we estimate the model using Japanese data slightly different from Koeda (2015).
16. Iwata and Fueda-Samikawa (2013) extensively examine the effects of the QQE on financial markets, including the analysis following Chadha *et al.* (2013). BOJ (2015) assesses the effects of the QQE on Japan's economic and financial developments during the two years since the introduction of the policy.
17. Using the U.S. flow of funds data, Carpenter *et al.* (2014) identify the types of investors that are selling to the Federal Reserve and investigate their portfolio adjustments after these sales, without examining the effects on interest rates. Saito and Hogen (2014) conduct similar analyses using Japanese data. Using a comprehensive dataset on the sovereign investor base for 45 advanced and emerging market economies, Jaramillo and Zhang (2013) examine the effects on bond yields of "real money investors," which include institutional investors as well as national and foreign central banks.
18. The fixed-rate JGBs cover the majority of the government debt in Japan (around 75 percent of the central government bonds and 60 percent of the general government gross debt as of 2013). We do not use data on floating-rate bonds, inflation-linked bonds, financial bills, Treasury bills (T-bills), or any other discount bonds in our analyses. Also, we exclude directly underwritten bonds and focus on marketable bonds.
19. To avoid an endogeneity problem in our analyses, we use the amount outstanding data mainly in face-value terms rather than market-value terms.

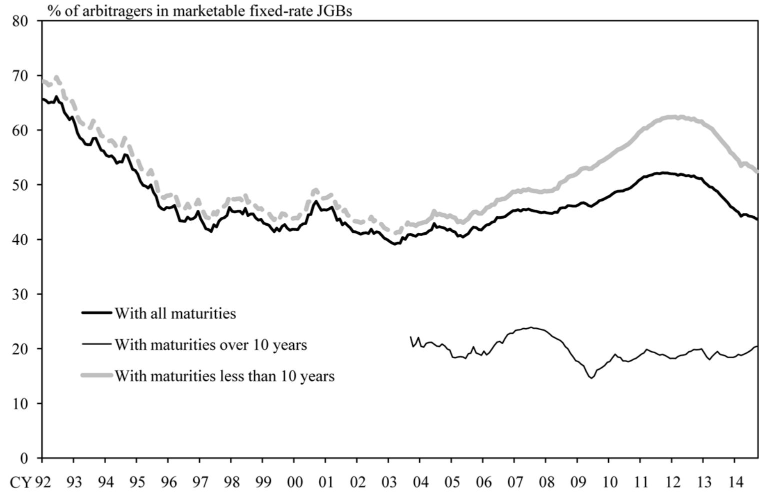
Figure 1 Amount Outstanding of Marketable Fixed-Rate JGBs Held by Each Sector/GDP



Note: Seasonally adjusted data. For details on data sources and the data calculation method, see Appendix 1.

shows those held by four sectors: the BOJ, insurance companies, pension funds, and "arbitragers." Throughout the paper, we regard insurance companies and pension funds as well as the BOJ as "preferred-habitat investors," because these private investors clearly prefer to hold long-term bonds to match their long-duration liabilities at least in our sample period as shown below. Accordingly, we define "arbitragers" as all JGB holders excluding the BOJ, insurance companies, and pension funds. Specifically, domestic banks, households, and foreigners are included in the arbitragers. The share of arbitragers in the total amount outstanding of fixed-rate JGBs has been within the range of 40 to 50 percent since the mid-1990s as shown in Figure 2.

Figure 2 Share of Arbitragers in Marketable Fixed-Rate JGBs



Note: Seasonally adjusted data. For details on data sources and the data calculation method, see Appendix 1.

As shown in Figure 1, the ratio of the amount outstanding of JGBs held by the arbitragers to GDP increased generally in parallel with the total debt-to-GDP ratio before the BOJ introduced the QQE in April 2013. However, it started to fall as the BOJ accelerated the pace of its JGB purchases as part of the QQE.[20] This movement in the arbitragers' debt-to-GDP ratio represents the movement in volume of the net supply of JGBs, which is defined as the issuance (supply) by the government minus the demand by the preferred-habitat investors, and it is used as a net supply measure in our analyses.

Figure 3 shows the maturity structures of JGBs held by the arbitragers, private preferred-habitat investors (insurance companies and pension funds),[21] and the BOJ from 2004 to the end of our sample period (September 2014). While the arbitragers' maturity structure has been relatively stable, the maturity structures of the preferred-habitat investors (both private and the BOJ) have changed remarkably. Private preferred-habitat investors, especially life insurance companies, steadily increased their holdings of JGBs with maturities over 10 years. They have needed to match the duration of assets to the long duration of their liabilities under regulations and accounting standards that force them to reduce their holdings of risky assets such as

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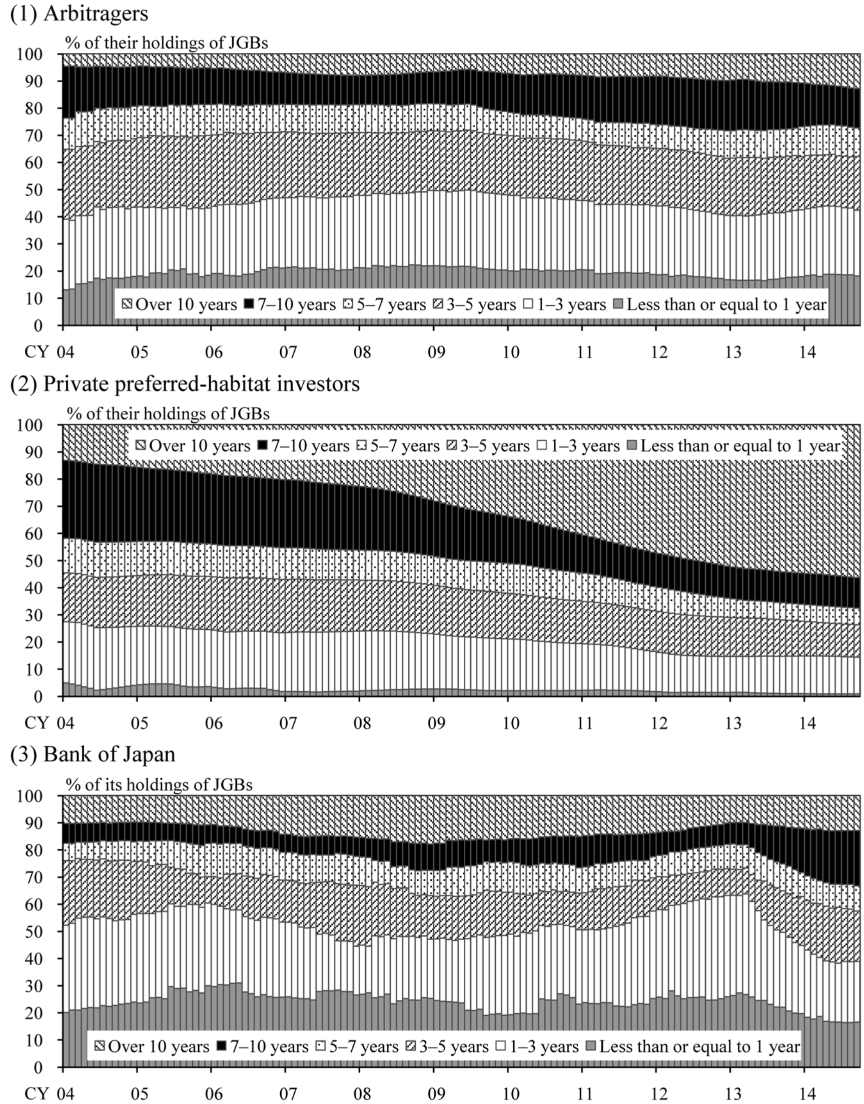
20. Before the introduction of the QQE, the BOJ gradually increased its purchase of JGBs as part of its monetary easing measures after the global financial crisis. Concurrently, pension funds have decumulated their asset holdings including JGBs to finance their rising pension payments since around 2008, and this has been an increasing factor in the arbitragers' debt-to-GDP ratio (Figure 1) and their share (Figure 2).

21. The "private" preferred-habitat investors actually include Japan Post Insurance (privatized in 2007) and public pension funds. See Appendix 1 for details.

Figure 3　Maturity Structure of JGBs Held by Each Sector



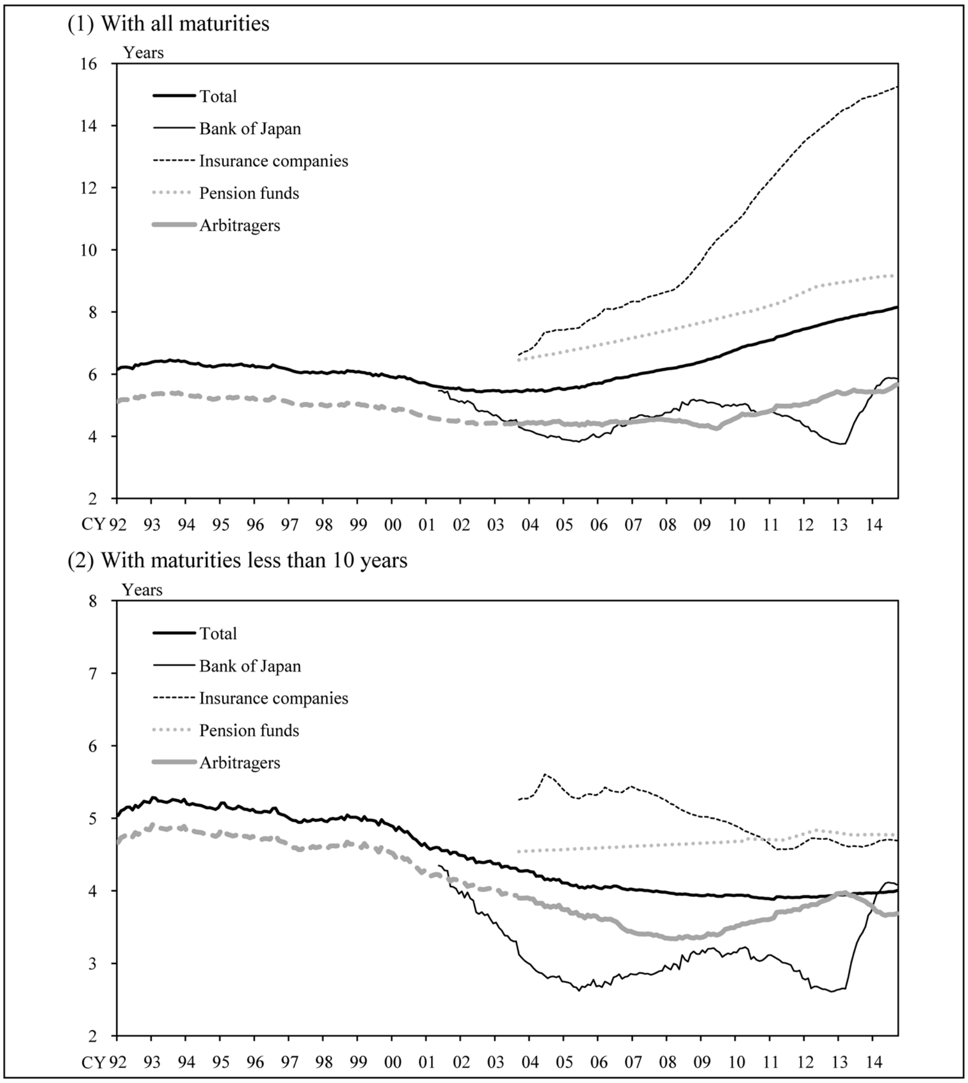
Note: Seasonally adjusted data. For details on data sources and the data calculation method, see Appendix 1.

stocks.[22] The BOJ has increased its holding of JGBs with maturities between 3 years and 10 years since the implementation of the QQE, while it has reduced the share of those with maturities less than 3 years in its JGB portfolio.

The average maturities of JGBs held by each sector as well as the total average

22. In addition to the prolonged low profitability of stocks since the bursting of the bubble economy in the 1990s, mark-to-market accounting (introduced in 2000) and solvency margin regulations (revised in 2012, when the risk coefficient for changes in domestic stock prices was raised from 10 percent to 20 percent) intensified Japanese life insurers' stance of reducing risks associated with stockholdings. See, for instance, Severinson and Yermo (2012) and Kan *et al.* (2013).

Figure 4 Average Maturities of JGBs Held by Each Sector

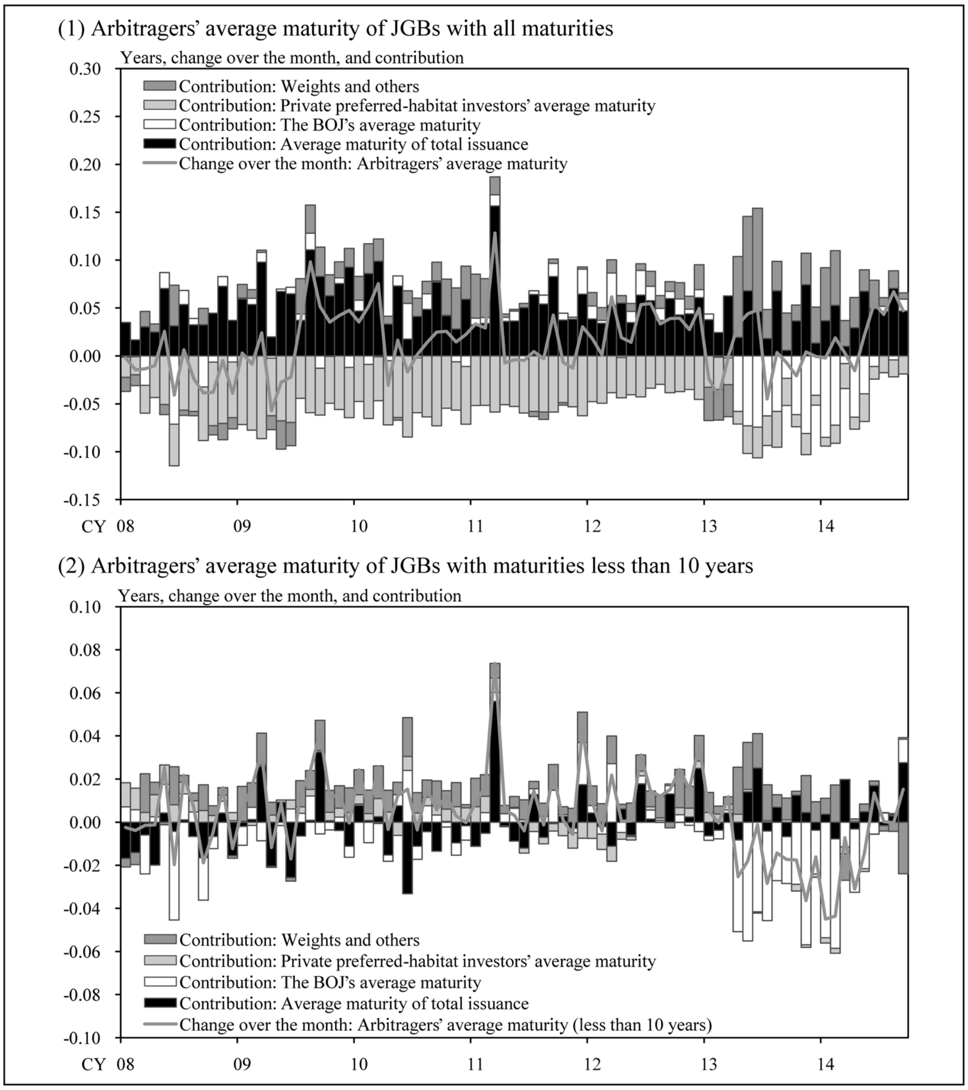


Note: Seasonally adjusted data. For details on data sources and the data calculation method, see Appendix 1.

maturities are shown in Figure 4. Insurance companies have sharply extended their average maturities while reducing the share of medium-term bonds in their JGB portfolio with maturities less than 10 years (shown in the lower panel). The BOJ has drastically extended its average maturity since the implementation of the QQE, especially in its JGB portfolio with maturities less than 10 years.[23] As a result, the arbitragers' average maturities, which had increased since around 2009, leveled off in their JGB

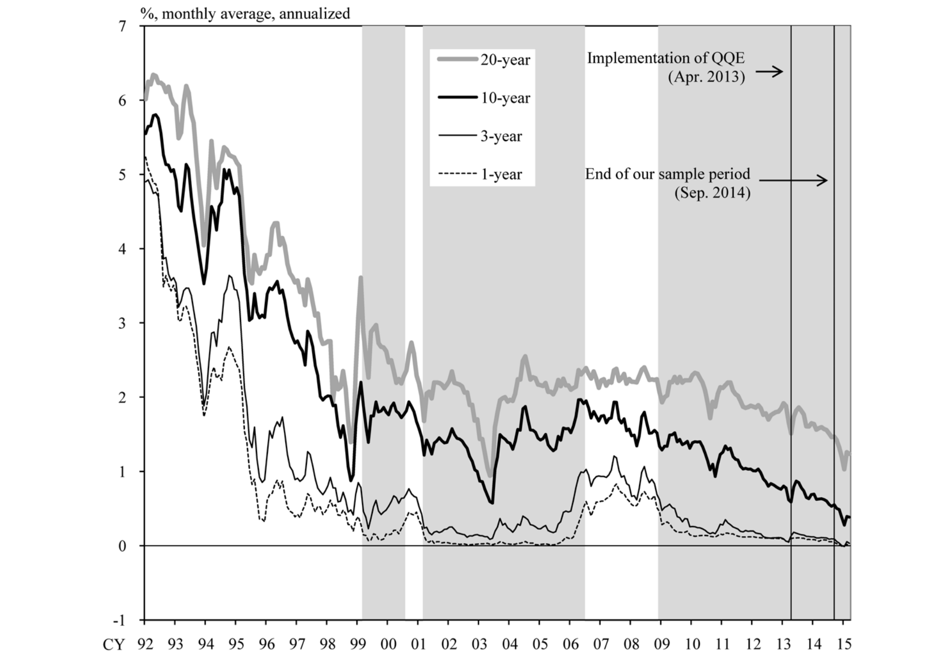
23. After the end of our sample period (not shown in Figure 4), the BOJ further extended its average maturity as it announced that "the average remaining maturity of the Bank's JGB purchases will be extended to about 7–10 years" on October 31, 2014 through an expansion of the QQE.

Figure 5 Decomposition of the Arbitragers' Average Maturities



portfolio with all maturities and started to decline in that with maturities less than 10 years. These average maturities of the arbitragers are used as key net supply measures for explaining term spreads in our analyses in Section III. In Figure 5, the changes in the arbitragers' average maturities since 2008 are decomposed into the changes in the average maturities of private preferred-habitat investors, the BOJ, and the total issuance by the government as well as the changes in weights among them. Before 2013, while the private preferred-habitat investors' maturity extension had been a decreasing factor in the arbitragers' average maturity, the maturity extension of the total issuance offset and overwhelmed it, which extended the arbitragers' average maturity as a result. After 2013, the BOJ's maturity extension was added as another decreasing factor, but it was offset by the effect of the increase in the BOJ's share that had a relatively shorter aver-

Figure 6 Zero-Coupon Yields



Note: The zero-coupon yield data are provided by Bloomberg. The shaded area indicates zero lower bound periods. The graph ends in March 2015.

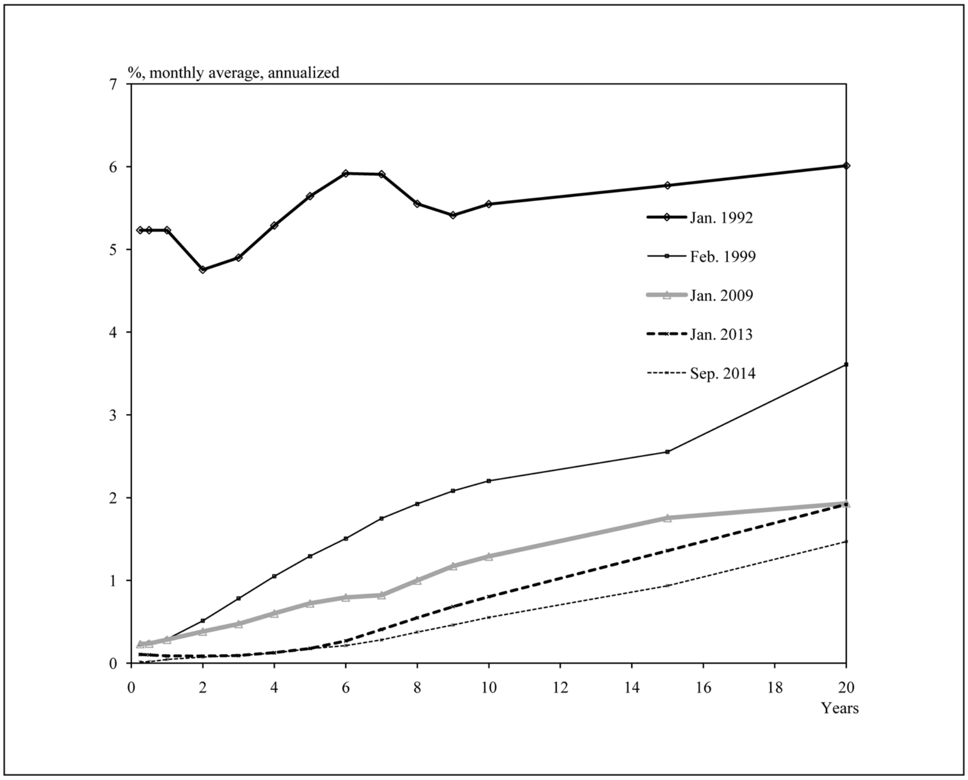
age maturity than the arbitragers as well as the effect of the maturity extension of the total issuance. Meanwhile, the BOJ's maturity extension overwhelmingly affected the arbitragers' average maturity in their JGB portfolio with maturities less than 10 years.

Finally, Figure 6 and Figure 7 show the data on JGB yields. Throughout the paper, we use the zero-coupon yields calculated by Bloomberg. As shown in Figure 6, the 1-year and 3-year yields have been very low since the zero interest rate policy was first implemented in 1999. The 10-year and 20-year yields have steadily declined since the mid-2000s, while the pace of decline in the 10-year yield has been faster than in the 20-year yield. In Figure 7, it can be clearly seen that the yield curve for relatively shorter maturities flattened between 2009 and 2013, and then the yield curve for maturities longer than 6 years flattened after 2013 as a result of the QQE.

III. Regression Approach

In this section, we examine the effects of the net supply of JGBs on the term structure of interest rates and the risk premium on long-term bonds using a regression approach. As the dependent variables representing the term structure, we focus on two measures of term spreads: the difference between zero-coupon yields on 10-year and 3-year JGBs

Figure 7 Yield Curves



Note: The zero-coupon yield data are provided by Bloomberg.

and the difference between the yields on 20-year and 10-year JGBs. The corresponding net supply measures we mainly use are the average maturity of JGBs with maturities less than 10 years held by the arbitragers (defined in Section II) and the average maturity of JGBs with all maturities held by the arbitragers, respectively. As the dependent variables representing the risk premium, we use the 1-year and 3-year holding-period excess returns of 10-year JGBs. The corresponding net supply measures are the ratio of JGBs with maturities less than 10 years held by the arbitragers to nominal GDP as well as the average maturity of those JGBs held by the arbitragers. Unless otherwise mentioned, we use monthly data from January 1992 to September 2014.

Our regression approach is closely related to Greenwood and Vayanos (2014, GV henceforth). However, besides the difference between gross and net supply measures mentioned in Introduction, our choices of dependent variables and net supply measures differ slightly from GV. They use long-term yields or returns rather than term spreads or excess returns as dependent variables while using a short-term (1-year) yield as a control variable (or alternatively, they use the spreads between long-term and short-term yields as dependent variables). We do not use the 1-year yield either for term spreads or as a control variable because it was very low and moved little due to the zero interest rate policy for a significant part of our sample period. Regarding the net supply measures, we do not use the maturity-weighted debt-to-GDP ratio, which is GV's main supply measure, because its movement was predominantly driven by the strong upward trend of the debt-to-GDP ratio in Japan. While the maturity-weighted debt-to-GDP ratio can be decomposed into the average maturity and the debt-to-GDP ratio, we use only the average maturity for the equations of term spreads and both the average maturity and the debt-to-GDP ratio separately for the equations of excess returns. Besides GV, we follow Laubach (2009) in controlling the economic fundamentals for the term spreads and Cochrane and Piazzesi (2005) in using their proposed return-forecasting factor as an explanatory variable for the holding-period excess returns.

There are mainly two econometric issues that need to be addressed in our regression approach. First, some of our net supply measures and control variables have a unit root, while the null hypotheses of a unit root are rejected for our dependent variables except for the 10-year/3-year yield spread. For the 10-year/3-year spread equation, we identify a single cointegrating relationship and estimate it by the fully-modified ordinary least squares (FM-OLS) method. For the other equations, we report level-regression results estimated by OLS with *t*-statistics using Newey and West (1987) standard errors to deal with serially correlated errors. Second, the arbitragers' average maturities could be endogenous and affected by (as well as affecting) the term spreads. We calculate average maturities mainly in face-value terms rather than market-value terms to avoid this endogeneity problem, but the concern still remains even with the face-value average maturities. In general, a widening of term spreads is likely to induce the government to shift the issuance of its debt toward shorter maturities, while it induces private preferred-habitat investors to hold government bonds with longer maturities. Both of these tendencies generate a negative relationship between term spreads and the arbitragers' average maturities and thus would bias our regression outcomes toward smaller net supply effects. We address this issue by taking the instrumental variable estimation approach.

In what follows, we present the estimation specifications and results for the 10-year/3-year spread, the 20-year/10-year spread, and the holding-period excess returns, in turn.

A. Regression of the 10-year/3-year Yield Spread

Our baseline specification of the regression equation of the 10-year/3-year yield spread, $y_t^{3,10}$, is expressed as follows:

$$y_t^{3,10} = \beta + \beta_x x_t^{10} + \beta_1 z_{1t} + \beta_3 z_{3t} + u_t, \tag{1}$$

where x_t^{10} is the average maturity of JGBs with maturities less than 10 years held by the arbitragers, z_{1t} is the 10-year/3-year spread of nominal GDP trend growth, and z_{3t} is the equity premium. β, β_x, β_1, and β_3 are parameters to be estimated and u_t is the error term. Following Laubach (2009), we try to control the expected trend growth in real consumption, the expected inflation, and the time variation in risk aversion as fundamental determinants of nominal long-term interest rates.[24] Due to unavailability

24. In the Ramsey model of optimal growth with a representative household with CES utility, the real rate of return on capital net of depreciation is determined by the net growth rate of per capita consumption, the intertemporal elasticity of substitution, and households' rate of time preference. For details, see Laubach

Table 1 Unit Root and Cointegration Tests

(1) Unit Root Test

Series	Augmented Dickey–Fuller test statistic	
	Level	First difference
10-year/3-year JGB zero-coupon yield spread	–2.261	–13.931 ***
20-year/10-year JGB zero-coupon yield spread	–4.434 ***	—
Arbitragers' average maturity (less than 10 years), IV	–1.740	–7.112 ***
Arbitragers' average maturity, IV	0.510	–4.515 ***
Nominal GDP trend-growth spread (10-year/3-year)	–2.170	–3.356 **
Nominal GDP trend-growth spread (20-year/10-year)	–0.937	–2.105 **
Equity premium	–1.438	–6.988 ***
Ex-post 1-year holding excess return	–2.205 **	—
Ex-post 3-year holding excess return	–2.030 **	—
Ratio of JGBs (less than 10 years) held by the arbitragers to nominal GDP	–0.168	–2.907 **
Return-forecasting factor	–2.538	–13.006 ***

(2) Johansen's Cointegration Test

Series: 10-year/3-year JGB zero-coupon yield spread, arbitragers' average maturity (less than 10 years, original), nominal GDP trend growth spread (10-year/3-year) equity premium

Null hypothesis	Trace statistic
None	64.835 ***
At most 1	27.751

Series: 10-year/3-year JGB zero-coupon yield spread, arbitragers' average maturity (less than 10 years, IV), nominal GDP trend growth spread (10-year/3-year) equity premium

Null hypothesis	Trace statistic
None	70.882 ***
At most 1	28.600

Note: *** and ** indicate statistical significance at the 1 percent and 5 percent levels, respectively.

of data on expected long-term trend growth and inflation for our entire sample period, we use a backward moving average with triangular weights of *nominal* GDP growth as a proxy for the combination of expected real trend growth and inflation. The spread is simply calculated as the difference between the 10-year moving average and the 3-year moving average. For risk aversion, we use an equity premium, which is a measure of expected excess returns of risky over risk-free assets, as a proxy.[25] We do not control the short-term deviations from the natural rate of interest caused by nominal rigidities as well as the counter-cyclical monetary policy rule, which would provide little information in addition to the nominal trend-growth spread for the long-term yield spread. Other candidates for control variables are discussed below.

As summarized in Table 1, according to the Dickey–Fuller test, the null hypothesis of a unit root for all the above dependent and explanatory variables cannot be rejected. We then conduct Johansen's cointegration test and confirm the existence of a single cointegrating relationship in the above set of variables.

Before estimating the identified cointegrating relationship of equation (1), we conduct the first-stage regression of the arbitragers' average maturity on instrumental variables to deal with the possible endogeneity problem mentioned above. Following GV, we use the debt-to-GDP ratio, which is driven mainly by the accumulation of past deficits rather than the current market conditions, as an instrumental variable. Specifically, the one-month lagged ratio of JGBs (less than 10 years) held by the preferred-

(2009).

25. Following Laubach (2009) and Kameda (2014), we use the equity premium based on the dividend yield as a proxy for risk aversion. More specifically, our measure is computed as the ratio of the dividend component of national income (SNA data) to the sum of the market value of "shares and other equities" (flow of funds data) held by households, minus the 10-year JGB yields. For this equity premium as well as GDP, we linearly interpolate the original quarterly data to obtain monthly data.

Table 2 Baseline Regression of 10-year/3-year Yield Spread

(1) First-Stage Regression

Dependent variable: Arbitragers' average maturity (less than 10 years), original

Ratio of JGBs (less than 10 years) held by the preferred-habitat investors to nominal GDP (-1)	–6.009 *** [–15.47]
Square of the ratio of JGBs (less than 10 years) held by the preferred-habitat investors to nominal GDP (-1)	4.757 *** [6.35]
Constant	5.289 *** [128.21]

Sample period	1992M01–2014M09
Estimation method	OLS
Adjusted R-squared	0.898

(2) Second-Stage Regression

Dependent variable: 10-year/3-year JGB zero-coupon yield spread

Arbitragers' average maturity (less than 10 years), IV	0.222 ** [2.12]
Nominal GDP trend-growth spread (10-year/3-year)	16.717 *** [4.48]
Equity premium	–0.101 ** [–2.11]
Constant	0.332 [0.65]

Sample period	1992M03–2014M09
Estimation method	FM-OLS
Adjusted R-squared	0.633

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

habitat investors (including the BOJ) to nominal GDP and its square are used here.[26] The results of this first-stage regression are summarized in the left part of Table 2. We also confirm the existence of a single cointegrating relationship in the obtained fitted values with the other explanatory variables and dependent variable in equation (1).

As the second-stage regression, we estimate the cointegrating relationship of equation (1) by the FM-OLS method to obtain asymptotically unbiased estimates.[27] The results are summarized in the right part of Table 2. The estimated coefficient on the arbitragers' average maturity is positive and significant at the 5 percent critical level. This implies that the net supply of relatively longer-term (but less than 10 years) JGBs tends to widen the 10-year/3-year spread, which is consistent with the preferred-habitat theory. The nominal trend-growth spread also has a significantly positive relationship with the long-term yield spread, which is consistent with neoclassical growth theory. The coefficient on the equity premium is significantly negative, which implies that an increase in households' (final investors') risk aversion tended to shift their demand from equities to 10-year JGBs rather than to 3-year JGBs and tighten the 10-year/3-year spread in our sample period. Figure 8 shows the decomposition of the fitted values (adjusted by the constant term) into three explanatory variables, from which we can see that the decrease in the arbitragers' average maturity together with a shrinking trend-growth spread have contributed to tightening of the 10-year/3-year spread in recent years.

Table 3 summarizes the results from the specifications using alternative measures of the net supply of JGBs, while the two control variables are the same as in the baseline specification. We could not identify a single cointegrating relationship in some of these specifications, but we report level-regression results estimated by OLS with *t*-

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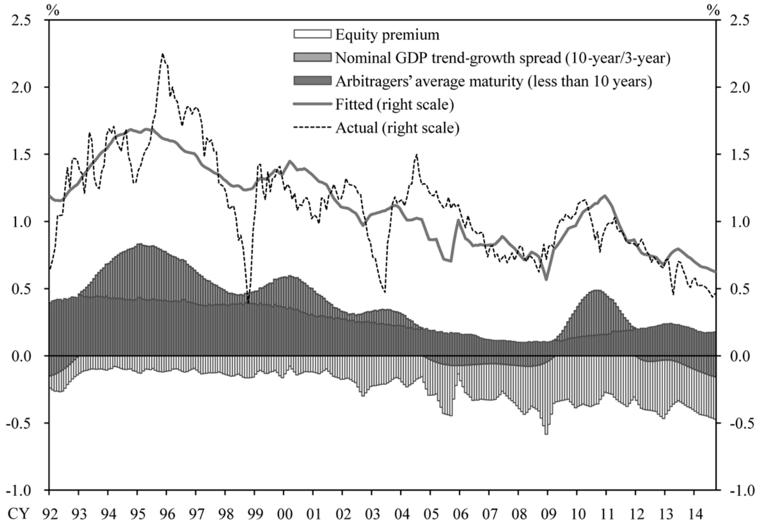
26. We use the preferred-habitat investors' debt-to-GDP ratio as an instrument because it is more exogenous (in terms of the *J*-statistics) with respect to the term spread and its upward trend is less steep than the total debt-to-GDP ratio. We have tried many alternative sets of instrumental variables and confirmed that our main results in our regression approach are generally robust.
27. In the FM-OLS estimation, we estimate the long-run covariance matrices in the cointegrating equation and stochastic regressors innovations using a kernel approach with a Bartlett kernel and Newey–West fixed bandwidth.

Figure 8 Decomposition of the 10-year/3-year Yield Spread



Note: The constant term is adjusted. For details on the results of the estimation, see Table 2. The arbitragers' average maturity in the above figure is the original (not instrumented) data.

statistics using Newey and West (1987) standard errors in Table 3. The column titled "baseline" shows the OLS estimation result of the baseline specification, in which the estimated coefficients are largely unchanged from the FM-OLS estimation result shown in Table 2. Column (i) shows the result using the original (not instrumented) data of the arbitragers' average maturity. The estimated coefficient on the average maturity is insignificant and smaller than that in the baseline result, reflecting the downward bias due to the endogeneity problem mentioned above.[28] In column (ii), the change in the ratio of JGBs (less than 10 years) held by the arbitragers to nominal GDP, smoothed by the Hodrick–Prescott filter, is included as an additional net supply measure that captures the movements in volume of the net supply. The estimated coefficient on this measure is positive but insignificant, while the signs and significance of the average maturity and control variables are maintained. In column (iii), we use the share of JGBs with maturities in excess of 3 years held by the arbitragers in their JGBs with maturities less than 10 years (the share of maturities of 3 to 10 years within those less than 10 years) as an alternative net supply measure.[29] The estimated coefficient on this measure is positive but insignificant.

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28. The estimated coefficient on the original average maturity is significant when we use the usual (not Newey–West) standard error (the t-statistic is 2.24). Meanwhile, the same coefficient estimated by FM-OLS is even smaller (0.08) and insignificant.
29. Modigliani and Sutch (1967) point out that the average maturity fails to distinguish changes occurring at the short end of the maturity range from those at the long end. They use in their analysis the proportion of the total volume of outstanding debt accounted for by issues falling within a specified maturity range as an alternative measure of maturity composition.

Table 3 Regression of 10-year/3-year Yield Spread (Alternative 1)

Dependent variable: 10-year/3-year JGB zero-coupon yield spread

	Baseline	Alternative net supply measures			
		(i)	(ii)	(iii)	(iv)
Arbitragers' average maturity (less than 10 years), IV	0.213 ** [2.35]		0.227 ** [2.46]		
Arbitragers' average maturity (less than 10 years), original		0.111 [1.27]			
Change in the ratio of JGBs (less than 10 years) held by the arbitragers to GDP, smoothed			7.619 [0.57]		
Share of 3- to 10-year JGBs held by arbitragers within less than 10 years				0.667 [0.86]	
BOJ's average maturity (less than 10 years)					–0.158 * [–1.79]
Arbitragers' and BOJ's average maturity (less than 10 years)					0.224 * [1.90]
Nominal GDP trend-growth spread (10-year/3-year)	17.273 *** [4.70]	18.528 *** [4.53]	16.583 *** [3.84]	18.804 *** [4.59]	18.620 *** [4.59]
Equity premium	–0.091 ** [–2.57]	–0.122 *** [–3.43]	–0.095 *** [–2.64]	–0.125 *** [–2.84]	–0.128 *** [–3.64]
Constant	0.338 [0.78]	0.811 ** [1.99]	0.279 [0.64]	0.885 * [1.73]	0.922 ** [2.15]
Sample period	1992M02-2014M09	1992M01-2014M09	1992M02-2014M09	1992M01-2014M09	1992M01-2014M09
Estimation method	OLS	OLS	OLS	OLS	OLS
Adjusted R-squared	0.630	0.612	0.630	0.609	0.615

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

Column (iv) shows the result using the average maturities of JGBs with maturities less than 10 years held by the BOJ and those issued by the government minus the holdings of private preferred-habitat investors separately (the latter equals the holdings of the arbitragers and the BOJ). The estimated coefficient on the BOJ's average maturity is significantly negative, which is consistent with the preferred-habitat theory. The absolute value of the estimated coefficient can be regarded as relatively large compared with the estimated coefficient on the other component of the net supply, considering that the BOJ's share in JGBs with maturities less than 10 years has been around 15 percent on average since the data became available in 2001 (while the share of the arbitragers plus the BOJ has been around 65 percent). This result may imply some heterogeneity between the BOJ and other (private) preferred-habitat investors and may also reflect the downward bias caused by the endogenous interaction between the term spread and the average maturities of the government and private preferred-habitat investors as mentioned above.[30] In Section V, we calculate the effect of the BOJ's JGB purchases and maturity extension as part of the QQE based on this result as well as the baseline result and other approaches.

Table 4 summarizes the results using alternative sets of control variables, while the net supply measure is the same as in the baseline specification. Column (v) shows the result using the real trend-growth spread (based on real GDP) and the trend-inflation spread (based on the consumer price index) instead of the nominal trend-growth spread. For both variables, backward moving averages with triangular weights are used and the

30. Kuttner (2006) shows that the effect of changes in the Federal Reserve's holdings of long-term securities had a larger and statistically significant effect on the holding-period excess returns than those in the composition of privately held debt because the latter includes the potentially endogenous issuance of new Treasury securities.

Table 4 Regression of 10-year/3-year Yield Spread (Alternative 2)

Dependent variable: 10-year/3-year JGB zero-coupon yield spread

	Baseline	Alternative control variables		
		(v)	(vi)	(vii)
Arbitragers' average maturity (less than 10 years), IV	0.213 ** [2.35]	0.272 ** [2.33]	0.222 ** [2.06]	0.217 ** [2.25]
Nominal GDP trend-growth spread (10-year/3-year)	17.273 *** [4.70]		17.114 *** [4.32]	17.304 *** [4.73]
Real GDP trend-growth spread (10-year/3-year)		12.002 *** [2.99]		
Trend-inflation spread (10-year/3-year)		0.173 ** [2.38]		
Volatility of 10-year/3-year JGB zero-coupon yield spread			–0.074 [–0.18]	
10-year/3-year U.S. Treasury zero-coupon yield spread				–0.014 [–0.36]
Equity premium	–0.091 ** [–2.57]	–0.087 ** [–2.13]	–0.092 ** [–2.58]	–0.084 * [–1.86]
Constant	0.338 [0.78]	0.122 [0.22]	0.313 [0.67]	0.321 [0.70]
Estimation method	OLS	OLS	OLS	OLS
Adjusted R-squared	0.630	0.593	0.629	0.629

Note: Sample period: 1992M02-2014M09. ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

spread is calculated as the difference between the 10-year moving average and the 3-year moving average. The estimated coefficients on both variables are significantly positive, while the adjusted R-squared is slightly lower than in the baseline regression. In columns (vi) and (vii), the volatility of the 10-year/3-year JGB zero-coupon yield spread (dependent variable) and the corresponding spread of the U.S. Treasury Bonds are included, respectively, as an additional control variable. However, the estimated coefficients on these additional variables are insignificant with wrong signs.

Finally in Table 5, we show the results of sub-sample regressions using the baseline specification. In column (viii), the sample period is limited to the three sub-periods under the zero lower bound of the short-term nominal interest rate: from March 1999 to July 2000, from March 2001 to June 2006, and from December 2008 to September 2014 (the end of the sample period), which correspond to the "zero lower bound (ZLB) regime" in the term structure model in Section IV. In column (ix), the sample period is further limited to the last sub-period after December 2008. In both results, especially in column (ix), the estimated coefficients on the arbitragers' average maturity are larger than that in the baseline full-sample regression, which imply that the net supply effect on the 10-year/3-year spread was stronger in the zero lower bound periods including the recent period after the global financial crisis. A possible interpretation based on the preferred-habitat theory is that the arbitragers' increasing aversion to the duration risks strengthened the net supply effect through the deepened market segmentation in the zero lower bound periods. To consider the possibility of the time-varying effect of the net supply depending on the risk aversion, we estimate another specification in which the arbitragers' average maturity in the baseline specification is replaced with the product of the average maturity and the equity premium. The latter in the product is used as an independent variable representing the risk aversion and was generally higher in the

Table 5 Regression of 10-year/3-year Yield Spread (Alternative 3)

Dependent variable: 10-year/3-year JGB zero-coupon yield spread

	Baseline	Sub-sample regressions		Time-varying net supply
		(viii)	(ix)	(x)
Arbitragers' average maturity (less than 10 years), IV	0.213 ** [2.35]	0.333 *** [3.09]	0.843 *** [3.77]	
Product of the average maturity and the equity premium				0.122 * [1.71]
Nominal GDP trend-growth spread (10-year/3-year)	17.273 *** [4.70]	7.138 *** [2.77]	13.795 *** [3.66]	18.372 *** [4.98]
Equity premium	–0.091 ** [–2.57]	–0.075 *** [–2.94]	0.021 [0.63]	–0.562 ** [–2.35]
Constant	0.338 [0.78]	–0.117 [–0.26]	–2.375 *** [–3.22]	1.192 *** [11.62]
Sample period	1992M02-2014M09	1999M03-2000M07 2001M03-2006M06 2008M12-2014M09	2008M12-2014M09	1992M02-2014M09
Estimation method	OLS	OLS	OLS	OLS
Adjusted R-squared	0.630	0.430	0.726	0.621

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted t-statistics are presented in the brackets.

zero lower bound periods. According to the result shown in column (x), the estimated coefficient on this risk-weighted net supply measure is positive and significant, while the signs and significance of other explanatory variables are maintained. This result together with the baseline result implies that the net supply effect on the 10-year/3-year spread was stronger when the risk aversion was higher, which supports the above conjecture based on the preferred-habitat theory.

B. Regression of the 20-year/10-year Yield Spread

Our baseline specification of the regression equation of the 20-year/10-year yield spread, $y_t^{10,20}$, corresponds with that of the 10-year/3-year spread equation, and is expressed as follows:

$$y_t^{10,20} = \beta + \beta_x x_t + \beta_2 z_{2t} + \beta_3 z_{3t} + u_t, \tag{2}$$

where x_t is the average maturity of JGBs (with all maturities) held by arbitragers,[31] z_{2t} is the 20-year/10-year spread of nominal GDP trend growth, and z_{3t} is the equity premium. As summarized in the left part of Table 1, according to the Dickey–Fuller test, the null hypothesis of a unit root for the dependent variable is strongly rejected at the 1 percent critical level, while those for all explanatory variables cannot be rejected.[32] Accordingly, we report level-regression results estimated by OLS with t-statistics using Newey and West (1987) standard errors.

As in the case of the 10-year/3-year spread equations, we conduct two-stage least square regression to deal with the possible endogeneity problem. In the first-stage regression, the arbitragers' average maturity is regressed on the one-month lagged

31. We use the arbitragers' average maturity of JGBs with all maturities rather than those over 10 years because the issuance of JGBs with maturities over 10 years was very limited in the early part of our sample period (especially before 1998).
32. According to Johansen's cointegration test, we can identify a single cointegrating relationship in the three explanatory variables in equation (2).

Table 6 Baseline Regression of 20-year/10-year Yield Spread

(1) First-Stage Regression

Dependent variable: Arbitragers' average maturity, original

Ratio of JGBs held by the preferred-habitat investors to nominal GDP (-1)	–5.315 *** [–35.86]
Square of the ratio of JGBs held by the preferred-habitat investors to nominal GDP (-1)	5.028 *** [20.53]
Share of JGBs over 10 years in total issuance (-1)	4.891 *** [14.35]
Constant	5.209 *** [111.71]
Sample period	1992M01-2014M09
Estimation method	OLS
Adjusted R-squared	0.911

(2) Second-Stage Regression

Dependent variable: 20-year/10-year JGB zero-coupon yield spread

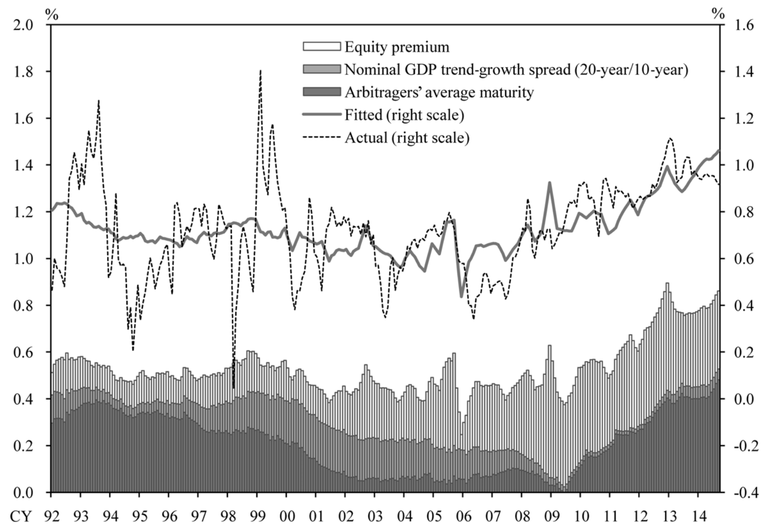
Arbitragers' average maturity, IV	0.348 *** [5.12]
Nominal GDP trend-growth spread (20-year/10-year)	6.202 * [1.85]
Equity premium	0.107 *** [4.28]
Constant	–1.290 *** [–3.28]
Sample period	1992M02-2014M09
Estimation method	OLS
Adjusted R-squared	0.279

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted t-statistics are presented in the brackets.

preferred-habitat investors' debt-to-GDP ratio, its square, and the share of JGBs with maturities over 10 years in the total issuance, as summarized in the left part of Table 6. In the second-stage regression, we estimate equation (2) using the fitted values of the first-stage regression. The results are summarized in the right part of Table 6. The estimated coefficient on the arbitragers' average maturity is positive and significant at the 1 percent critical level, which implies that the net supply of relatively longer-term JGBs tends to widen the 20-year/10-year spread as well as the 10-year/3-year spread, which is consistent with the preferred-habitat theory. The nominal trend-growth spread also has a significantly positive relationship with the super-long-term yield spread, consistently with neoclassical growth theory. The coefficient on the equity premium is significantly positive rather than negative, in contrast to the result for the 10-year/3-year spread. It may be naturally understood that an increase in investors' risk aversion tends to widen the 20-year/10-year spread because they become averse to the duration risk for this super-long-term spread. In our sample period, the increasingly risk-averse investors tended to shift their demand from equities to 10-year JGBs rather than to 20-year JGBs or 3-year JGBs. Figure 9 shows the decomposition of the fitted values (adjusted by the constant term) into three explanatory variables, from which we can see that the increase in the arbitragers' average maturity has contributed to the widening of the 20-year/10-year spread since around the recent financial crisis.

Table 7 summarizes the results from the specifications using alternative measures of the net supply of JGBs, while the two control variables are the same as in the baseline specification. The column titled "baseline" shows exactly the same result as in Table 6. Column (i) shows the result using the original (not instrumented) data of the arbitragers' average maturity. The estimated coefficient on the average maturity is slightly smaller than that obtained from the instrumental variable estimation, although the difference is much smaller than in the case of the 10-year/3-year spread equations. In column (ii), the change in the ratio of JGBs held by the arbitragers to nominal GDP, smoothed by the Hodrick–Prescott filter, is included as an additional net supply measure. The estimated coefficients on this measure as well as the average maturity of the net supply are both significantly positive. In column (iii), we use the share of JGBs with maturities in excess of 10 years held by the arbitragers in their total holdings of

Figure 9 Decomposition of the 20-year/10-year Yield Spread



Note: The constant term is adjusted. For details on the results of the estimation, see Table 6. The arbitragers' average maturity in the above figure is the original (not instrumented) data.

JGBs as an alternative net supply measure. The estimated coefficient on this measure is positive and significant at the 1 percent critical level.

Column (iv) shows the result using the average maturities of JGBs held by the BOJ, those held by private preferred-habitat investors, and those issued by the government, separately. The estimated coefficients on the average maturities of the BOJ and private preferred-habitat investors are significantly negative, which are consistent with the preferred-habitat theory. The magnitudes of these estimated coefficients generally correspond to their shares in the total outstanding of JGBs. (The BOJ's share has been around 15 percent on average since the data became available in 2001, while the private preferred-habitat investors' share has been around 40 percent.) On the one hand, there may be some downward bias caused by the endogeneity problem in the estimated coefficient on the private preferred-habitat investors' average maturity. On the other hand, since the BOJ's average maturity has been far below 10 years, its effect on the 20-year/10-year spread may be limited compared with the private preferred-habitat investors' average maturity that has increased to well over 10 years as shown in Figure 4 in Section II.

Table 8 summarizes the results using alternative sets of control variables, while the net supply measure is the same as in the baseline specification. Column (v) shows the result using the real trend-growth spread and the trend-inflation spread instead of the nominal trend growth spread. The estimated coefficients on the real trend-growth spread is significantly positive, while that on the trend-inflation spread is insignificant. In columns (vi) and (vii), the volatility of the 20-year/10-year JGB yield spread (depen-

Table 7 Regression of 20-year/10-year Yield Spread (Alternative 1)

Dependent variable: 20-year/10-year JGB zero-coupon yield spread

	Baseline	Alternative net supply measures			
		(i)	(ii)	(iii)	(iv)
Arbitragers' average maturity, IV	0.348 *** [5.12]		0.505 *** [4.05]		
Arbitragers' average maturity, original		0.314 *** [5.75]			
Change in the ratio of JGBs held by the arbitragers to GDP, smoothed			31.979 ** [2.04]		
Share of JGBs over 10 years held by arbitragers in their total holdings				4.905 *** [3.81]	
BOJ's average maturity					–0.081 *** [–3.02]
Private preferred-habitat investors' average maturity					–0.189 * [–1.91]
Average maturity of total issuance					0.673 *** [2.60]
Nominal GDP trend-growth spread (20-year/10-year)	6.202 * [1.85]	4.557 [1.38]	9.592 ** [2.17]	2.281 [0.60]	10.248 ** [2.05]
Equity premium	0.107 *** [4.28]	0.104 *** [4.28]	0.098 *** [4.12]	0.031 [1.19]	0.061 ** [2.17]
Constant	–1.290 *** [–3.28]	–1.093 *** [–3.37]	–2.155 *** [–3.05]	0.293 * [1.66]	–1.857 ** [–2.18]
Sample period	1992M02-2014M09	1992M01-2014M09	1992M02-2014M09	1992M01-2014M09	1992M01-2014M09
Estimation method	OLS	OLS	OLS	OLS	OLS
Adjusted R-squared	0.279	0.293	0.305	0.179	0.328

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

dent variable) and the corresponding spread of the U.S. Treasury Bonds are included, respectively, as an additional control variable. While the estimated coefficient on the former is insignificant, that on the latter is strongly significant and positive, in contrast to the result for the 10-year/3-year spread, and greatly improves the adjusted R-squared from the baseline result.

Finally in Table 9, we show the results of sub-sample regressions using the baseline specification. As in the results for the 10-year/3-year spread, column (viii) shows the result when the sample period is limited to the three sub-periods under the zero lower bound of the short-term nominal interest rate, and column (ix) shows the result when the sample period is further limited to the last sub-period after December 2008. In both results, the estimated coefficients on the arbitragers' average maturity are smaller than that in the full sample period, in contrast to the results for the 10-year/3-year spread. Meanwhile, column (x) shows that the estimated coefficient on the risk-weighted net supply measure, the product of the arbitragers' average maturity and the equity premium, is significantly positive, as in the result for the 10-year/3-year spread. The latter result implies that the net supply effect on the 20-year/10-year spread was stronger when the arbitragers' risk aversion was higher, while the former implies that the effect was *not* stronger in the zero lower bound periods when the risk aversion represented by the equity premium was higher. One possible factor that could reconcile these seemingly conflicting results is the share of JGBs held by the arbitragers in the total amount outstanding, which declined in the early 1990s and then dropped again after the recent financial crisis (especially their share in JGBs with maturities over 10 years) as shown in Figure 2 in Section II. When the presence of the arbitragers is small, the net

Table 8 Regression of 20-year/10-year Yield Spread (Alternative 2)

Dependent variable: 20-year/10-year JGB zero-coupon yield spread

	Baseline	Alternative control variables		
		(v)	(vi)	(vii)
Arbitragers' average maturity, IV	0.348 *** [5.12]	0.305 *** [5.45]	0.350 *** [5.38]	0.315 *** [4.99]
Nominal GDP trend-growth spread (20-year/10-year)	6.202 * [1.85]		6.266 ** [2.00]	6.541 ** [2.13]
Real GDP trend-growth spread (20-year/10-year)		5.640 * [1.71]		
Trend-inflation spread (20-year/10-year)		0.023 [0.54]		
Volatility of 20-year/10-year JGB zero-coupon yield spread			–0.036 [–0.07]	
20-year/10-year U.S. Treasury zero-coupon yield spread				0.317 *** [3.37]
Equity premium	0.107 *** [4.28]	0.102 *** [3.71]	0.106 *** [3.62]	0.071 *** [3.06]
Constant	–1.290 *** [–3.28]	–1.036 *** [–3.13]	–1.294 *** [–3.39]	–1.176 *** [–3.31]
Estimation method	OLS	OLS	OLS	OLS
Adjusted R-squared	0.279	0.270	0.276	0.372

Note: Sample period: 1992M02-2014M09. ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

Table 9 Regression of 20-year/10-year Yield Spread (Alternative 3)

Dependent variable: 20-year/10-year JGB zero-coupon yield spread

	Baseline	Sub-sample regressions		Time-varying net supply
		(viii)	(ix)	(x)
Arbitragers' average maturity, IV	0.348 *** [5.12]	0.314 *** [8.38]	0.317 *** [2.68]	
Product of the average maturity and the equity premium				0.124 *** [5.49]
Nominal GDP trend-growth spread (20-year/10-year)	6.202 * [1.85]	0.161 [0.08]	–13.875 [–1.05]	2.490 [0.82]
Equity premium	0.107 *** [4.28]	0.036 * [1.80]	0.015 [0.34]	–0.526 *** [–5.03]
Constant	–1.290 *** [–3.28]	–0.809 *** [–3.64]	–0.685 [–1.13]	0.537 *** [5.18]
Sample period	1992M02-2014M09	1999M03-2000M07 2001M03-2006M06 2008M12-2014M09	2008M12-2014M09	1992M02-2014M09
Estimation method	OLS	OLS	OLS	OLS
Adjusted R-squared	0.279	0.517	0.432	0.273

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets.

supply effect would not change significantly in response to changes in the arbitragers' risk aversion. Another possibility is that long-term (as well as short-term) interest rates become less responsive to any factors, including the net supply shocks, when the short-term interest rate is constrained by the zero lower bound. We check this possibility in Section IV using a term structure model that explicitly considers the zero lower bound regime as well as the normal regime.

C. Regression of Holding-Period Excess Returns

Next, we present the estimation specifications and results for the holding-period excess returns. Our specifications of the 1-year and 3-year holding excess returns of 10-year JGBs, $rx^{10}_{t+1,12}$ and $rx^{10}_{t+1,36}$, respectively, are expressed as follows:

$$\begin{aligned} rx^{10}_{t+1,12} &\equiv (10 \times y^{10}_t - 9 \times y^{9}_{t+12}) - y^{1}_t \\ &= \beta_{01} + \beta_x x^{10}_t + \beta_{xx} xx^{10}_t + \beta_4 z_{4t} + u_{1t}, \end{aligned} \tag{3}$$

$$\begin{aligned} rx^{10}_{t+1,36} &\equiv \frac{(10 \times y^{10}_t - 7 \times y^{7}_{t+36})}{3} - y^{3}_t \\ &= \beta_{03} + \beta_x x^{10}_t + \beta_{xx} xx^{10}_t + \beta_4 z_{4t} + u_{3t}, \end{aligned} \tag{4}$$

where y^n_t is the yield on n-year JGBs in period (month) t, x^{10}_t is the average maturity of JGBs with maturities less than 10 years held by the arbitragers, xx^{10}_t is the ratio of those JGBs held by the arbitragers to nominal GDP, and z_{4t} is the return-forecasting factor proposed by Cochrane and Piazzesi (2005).[33] We use xx^{10}_t as well as x^{10}_t as net supply measures to capture the volume as well as the maturity structure of the net supply. As mentioned in the beginning of this section, the product of these two measures corresponds to the maturity-weighted debt-to-GDP ratio, which is the main supply measure used by GV. Regarding the return-forecasting factor, it is constructed from the fitted value of a regression of $(1/4)\sum_{n=2}^{5} rx^n_{t+1,12}$ on a constant and the 1- through 5-year forward rates. As we use the realized (ex-post) holding-period excess returns as dependent variables, the end of the sample period is September 2013 for the 1-year holding excess return equation (3), and is September 2011 for that of the 3-year holding excess return equation (4).

As summarized in the left part of Table 1, according to the Dickey–Fuller test, the null hypotheses of a unit root for the dependent variables are rejected at the 5 percent critical level, while those for all explanatory variables cannot be rejected.[34] Accordingly, we report level-regression results estimated by OLS with t-statistics using Newey and West (1987) standard errors. The results are summarized in Table 10. Regarding the 1-year holding excess return, the estimated coefficients on the arbitragers' average maturity and their debt-to-GDP ratio are both positive but only the latter coefficient is significant, which implies that a larger volume of the net supply of JGBs tends to increase the risk premium on the 10-year bond. When the sample period is limited to the three sub-periods under the zero lower bound of the short-term nominal interest rate, as shown in column (ii), the estimated coefficients on the two net supply measures are larger than those in the full sample period, although the coefficient on the average maturity is still insignificant. Regarding the 3-year holding excess return, as shown in column (v), the estimated coefficients on the two net supply measures are both positive and significant at the 1 percent critical level, which implies that a relatively longer-term

33. Kuttner (2006) uses Cochrane and Piazzesi's (2005) return-forecasting factor as a control variable in his regression of the holding-period excess returns on the maturity structure of the Federal Reserve's holding securities.
34. According to Johansen's cointegration test, we can identify a single cointegrating relationship in the three explanatory variables in equations (3) and (4).

Table 10 Regression of Holding-Period Excess Returns

	Dependent variable: ex-post 1-year holding excess return of 10-year JGBs				Dependent variable: ex-post 3-year holding excess return of 10-year JGBs			
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)
Arbitragers' average maturity (less than 10 years), original	1.202 [0.98]	2.697 [1.59]		2.854 ** [2.23]	2.056 *** [5.19]	2.809 *** [4.76]		2.777 *** [3.03]
Ratio of JGBs (less than 10 years) held by the arbitragers to nominal GDP	5.976 ** [2.06]	8.114 * [1.82]		14.053 *** [2.98]	6.170 *** [4.41]	8.250 *** [4.14]		9.028 ** [2.22]
Product of the average maturity, the debt/GDP, and the equity premium			0.215 [1.25]				0.045 [0.64]	
Wealth measure				–0.015 ** [–2.52]				–0.017 [–0.53]
Return-forecasting factor	218.480 *** [5.08]	242.678 * [1.89]	224.066 *** [5.16]	189.317 *** [3.73]	105.429 *** [10.07]	163.674 *** [4.04]	122.355 *** [11.82]	135.630 *** [9.11]
Constant	–5.246 [–0.98]	–11.879 [–1.43]	0.699 [0.57]	–12.948 ** [–2.19]	–8.115 *** [–4.27]	–12.112 *** [–4.05]	1.553 *** [4.22]	–11.786 *** [–2.67]
Sample period	1992M01-2013M09	1999M03-2000M07 2001M03-2006M06 2008M12-2013M09	1992M01-2013M09	1993M01-2013M09	1992M01-2011M09	1999M03-2000M07 2001M03-2006M06 2008M12-2011M09	1992M01-2011M09	1995M01-2011M09
Estimation method	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
Adjusted R-squared	0.334	0.109	0.330	0.293	0.742	0.491	0.690	0.709

Note: ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. Newey–West adjusted *t*-statistics are presented in the brackets. The wealth measure is defined as the product of the 1-year (or 3-year) lag of the arbitragers' maturity-weighted debt-to-GDP ratio (the product of the two net supply measures) and the 1-year (or 3-year) holding excess return.

and a larger volume of the net supply of JGBs tends to increase the risk premium on the 10-year bond. When the sample period is limited to the three sub-periods under the zero lower bound, as shown in column (vi), the estimated coefficients on the two net supply measures are larger than those in the full sample period, while the signs and significance are maintained. This implies that the net supply effect on the risk premium was larger in the zero lower bound periods.

We then consider the possibility of the time-varying effect of the net supply depending on the arbitragers' risk aversion by estimating two alternative specifications. First, as in the analysis of term spreads, the two net supply measures are replaced with the product of the two measures and also the equity premium. As shown in columns (iii) and (vii) for the 1-year and 3-year holding excess returns, respectively, the estimated coefficients on these risk-weighted net supply measures are both positive but insignificant. Second, following GV, the product of the two net supply measures and the measure of the change in the arbitragers' wealth during the holding periods is included as an explanatory variable, in addition to the two net supply measures. The wealth measure is defined as the product of the 1-year (or 3-year) lag of the arbitragers' maturity-weighted debt-to-GDP ratio (the product of the two net supply measures) and the 1-year (or 3-year) holding excess return. As shown in columns (iv) and (viii) for the 1-year and 3-year holding excess returns, respectively, the estimated coefficients on these wealth measures are both negative, although only the coefficient in the 1-year return equation is significant. This implies that the net supply effects on the risk premium are stronger when the arbitragers have lost their wealth and thus their risk aversion is higher.

Throughout this section, from our regression approach, we can conclude that the net supply effects on the term structure of interest rates and the risk premium on long-term bonds are statistically significant in most alternative specifications, which is consistent with the preferred-habitat theory. Moreover, many regression results imply that the effects are stronger when the short-term interest rate is constrained by the zero lower bound and the arbitragers' risk aversion is higher.

IV. Term Structure Model Approach

In this section, we report the results from a term structure model that incorporates preferred-habitat investors.[35] In the model, yields for all maturities are expressed as affine functions of three yield-curve factors. Moreover, we consider the zero lower bound (ZLB) regime as well as the normal regime, allowing the coefficients on the yield-curve factor dynamics and the prices of risk to differ between the two regimes. Using this model, we examine the effects of the model-implied net supply measures of JGBs on term spreads and holding-period excess returns as in our regression approach in Section III. We also use this model in Section V to derive the term premium and examine the effects of the BOJ's JGB purchases and maturity extension.

Here we just give a brief outline of the model of the normal regime. Details of the model including the ZLB regime are described in Appendix 2 and Koeda (2015).

There are two types of investors in the JGB market: arbitragers and preferred-habitat investors. The arbitragers choose a fraction, z_{nt}, of their JGB portfolio in each maturity n (from 1 to N) to maximize the mean-variance expected returns on the portfolio, as follows:

$$\max_{z_{1t},\cdots,z_{Nt}} \left\{ E(r_{t,t+1}|\mathbf{f}_t) - \frac{\gamma}{2}\,\mathrm{Var}(r_{t,t+1}|\mathbf{f}_t) \right\},$$

where $r_{t,t+1} = \Sigma_{n=1}^{N} z_{nt}\, r_{n,t,t+1}$ is the arbitragers' rate of return from period t to period $t+1$ on their JGB portfolio and $\mathbf{f}_t$ is a 3×1 vector of yield-curve factors assumed to follow a VAR (1) process: $\mathbf{f}_{t+1} = \mathbf{c} + \mathbf{\Phi}\mathbf{f}_t + \mathbf{\Sigma}\mathbf{u}_{t+1}$. The parameter γ captures the arbitragers' risk aversion. In the ZLB regime, arbitragers explicitly consider the possibility of a regime shift to the normal regime in their optimization problem, as described in Appendix 2.

On the other hand, the preferred-habitat investors demand JGBs with a particular maturity depending solely on the corresponding yield on that particular maturity: they demand more of the bonds when the yield on that maturity increases. Meanwhile, the government would also decide to issue its bonds with a maturity depending on the yield on that particular maturity: it issues fewer of the bonds when the yield on that maturity increases. Therefore, the net supply of the JGBs with maturity n, the issuance by the government minus preferred-habitat investors' demand, can be expressed as:

$$x_{nt} = \varsigma_{t,n} - \alpha_n\, y_{t,n},$$

for each n, where x_{nt} is the net supply of the JGBs with maturity n divided by the arbitragers' net wealth, $\varsigma_{t,n}$ is an affine function of $\mathbf{f}_t$, and the parameter α_n represents the sensitivity to the n-period log yield, $y_{t,n}$. Then, in the market equilibrium where

35. The model used in this paper is the same as in Koeda (2015), which is an extension of Hamilton and Wu (2012a). The model of the normal regime described in the text is the same as in Hamilton and Wu (2012a).

$z_{nt} = x_{nt}$ for all n, the n-period log bond prices, $p_{t,n}$, and log yields, $y_{t,n}$, can be expressed as an affine function of $\mathbf{f}_t$ as:

$$p_{t,n} = -n\, y_{t,n} = \bar{a}_n + \bar{\mathbf{b}}_n\, \mathbf{f}_t,$$

for all n, given that the risk-free one-period rate, $y_{t,1}$, is assumed to follow an affine function of $\mathbf{f}_t$.

From this model, we can derive the term spread between j- and k-period yields $(j > k)$, $y_{t,j} - y_{t,k}$, and the expected k-period holding excess return of j-period bonds, $E_t(rx^j_{t+1,k})$, as follows:

$$\begin{aligned} y_{t,j} - y_{t,k} = {} & \tilde{c}_{j,k} + \tilde{\mathbf{\Gamma}}_{j,k}\mathbf{f}_t \\ & + \gamma \left\{ \left(\frac{\bar{\mathbf{b}}_{j-1}}{j} - \frac{\bar{\mathbf{b}}_{k-1}}{k} \right) \mathbf{\Sigma}\mathbf{\Sigma}' \sum_{n=2}^{N} z_{nt} \bar{\mathbf{b}}'_{n-1} \right\}, \end{aligned} \tag{5}$$

$$E_t(rx^j_{t+1,k}) = c_{j,k} + \mathbf{\Phi}_{j,k}\mathbf{f}_t + \gamma \bar{\mathbf{b}}_{j-1} \mathbf{\Sigma}\mathbf{\Sigma}' \sum_{n=2}^{N} z_{nt} \bar{\mathbf{b}}'_{n-1}, \tag{6}$$

where $\tilde{c}_{j,k}$, $\tilde{\mathbf{\Gamma}}_{j,k}$, $c_{j,k}$, and $\mathbf{\Phi}_{j,k}$ are combinations of parameters in the model. The point here is that $z_{nt}(= x_{nt})$ is included in the third term on the right-hand side of equations (5) and (6).[36] We regard this term as the model-implied net supply measure in each equation.

We estimate the model using monthly data of the zero coupon yield (the same as in Section III) on 3-month, 3-year, 5-year, and 10-year JGBs. As in Koeda (2015), we apply Hamilton and Wu (2012b) minimum chi-squared estimation to estimate the term structure model coefficients.[37] Meanwhile, some deep parameters including γ representing the risk aversion are not estimated in the model. In this paper, we try to estimate γ in equations (5) and (6) by OLS, using the actual data of term spreads and ex-post holding-period excess returns as dependent variables, respectively.[38] We construct the model-implied net supply measures as explanatory variables for the above regressions, by plugging in the data on JGB's maturity structure described in Section II for z_{nt} in the third term on the right-hand side of equations (5) and (6). In these regressions, the sample period is from January 1992 to September 2014, as in the regression analysis in Section III. We distinguish between the normal and ZLB regimes, the latter of which consists of three sub-periods: from March 1999 to July 2000, from March 2001 to June 2006, and from December 2008 to September 2014,[39] by using dummy variables.

In what follows, we first check the model-implied relationship between the maturity structure (z_{nt}) and the net supply measures in term spreads and holding-period

36. As described in Appendix 2, the price of risk can be expressed as a function of z_t in this model.
37. We first estimate the term structure model coefficients under the normal regime using the data from February 1992 to February 1999 and then, given the normal-regime parameters, the term structure model coefficients under the ZLB regime are estimated using the data from December 2008 to September 2014.
38. In these regressions, we relax the model assumption by allowing the net supply of the JGBs to depend on not only bond yields but also other possible exogenous shocks.
39. The identification of the ZLB regime is based on Hayashi and Koeda (2014).

excess returns, based on Koeda (2015). Then given these net supply measures and the estimated yield-curve factors, we try to estimate γ in equations (5) and (6), and discuss the results including the size, significance, and difference between the normal and ZLB regimes.

A. Regression of the 10-year/3-year Yield Spread

First we examine the effects of the model-implied net supply measure on the 10-year/3-year spread. The net supply measure, the third term on the right-hand side of equation (5), increases when the average maturity of JGBs held by the arbitragers is extended. We check how the change in the maturity structure less than 10 years (z_{nt} for $n \leq 10$) is related to the net supply measure, using the data on JGB's maturity structure described in Section II for z_{nt}. Suppose a shock causes the arbitragers to extend their maturity by buying 10-year bonds and selling 1-year bonds by a 1 percent share of their holding JGBs with maturities less than 10 years (at the same time, the preferred-habitat investors shorten their maturity by selling 10-year bonds and buying 1-year bonds) under the normal regime. Then, the net supply measure widens the 10-year/3-year yield spread by about 0.23 basis point, given the estimated term structure model coefficients and assuming the parameter value of the arbitragers' risk aversion to be $\gamma = 100$.[40] If, on the other hand, the arbitragers shorten their maturity (the preferred-habitat investors extend their maturity) by carrying out the opposite transaction under the ZLB regime, the net supply measure tightens the 10-year/3-year yield spread by about 0.06 basis point. Note that these figures in terms of the 10-year/3-year yield spread are very small because we focus here on the effect of the net supply measure (the third term in equation (5)) and do not take account of the changes in the yield-curve factors (the second term).[41] The asymmetry between the two regimes reflects the model estimation results in which the yield-curve factors are more volatile and yields are more responsive to the yield-curve factors under the normal regime, as in Koeda (2015). Therefore, according to this model, the net supply effect on the 10-year/3-year spread is stronger under the normal regime than under the ZLB regime, unless the degree of the arbitragers' risk aversion changes significantly between the two regimes.

Then, given the model-implied net supply measure and yield-curve factors, we estimate γ in equation (5) by OLS. We also estimate the coefficients on the yield-curve factors under the two alternative assumptions: they are identical or different between the two regimes.[42] As summarized in Table 11, in both specifications with and without the coefficient dummies for the ZLB regime on the yield-curve factors, the estimated coefficient on the net supply measure, γ, is significantly positive. Since γ represents the arbitragers' risk aversion in this model, this result implies not only that the net supply effect is significantly positive but also that the effect is stronger when the arbitragers' risk aversion is higher. However, the size of the estimated γ, differs substantially between the two specifications. Table 11 also shows the result when the

40. Hamilton and Wu (2012a) calibrate this parameter to 100 based on their estimation results using U.S. data.
41. When we calculate the effects of the BOJ's JGB purchases and maturity extension in Section V, we take account of the changes in the yield-curve factor as well as the net supply measure.
42. The coefficients $\tilde{\boldsymbol{\Gamma}}_{j,k}$ in equation (5) are estimated in the term structure model and differ between the two regimes. However, we do not use these estimated coefficients but re-estimate them in the single-equation regression to improve the fit of the regression.

Table 11 Regression of 10-year/3-year Yield Spread from the Term Structure Model Approach

Dependent variable: 10-year/3-year JGB zero-coupon yield spread

	(i)	(ii)	(iii)
Model-implied net supply measure	124.919 * [1.82]	244.774 ** [2.58]	136.141 [0.88]
Yield-curve factor 1	0.145 *** [39.86]	0.092 *** [18.11]	0.096 *** [52.39]
Yield-curve factor 2	–0.018 *** [–4.73]	0.010 *** [2.97]	0.047 *** [57.67]
Yield-curve factor 3	–0.025 *** [–9.93]	–0.001 *** [–2.96]	–0.001 *** [–9.60]
ZLB dummy × yield-curve factor 1	–0.049 *** [–11.06]		
ZLB dummy × yield-curve factor 2	0.066 *** [15.67]		
ZLB dummy × yield-curve factor 3	0.025 *** [9.58]		
Constant	1.400 *** [45.39]	1.114 *** [33.05]	1.066 *** [47.94]
ZLB dummy	–0.334 *** [–14.31]		
Sample period	1992M01-2014M09	1992M01-2014M09	1999M03-2000M07 2001M03-2006M06 2008M12-2014M09
Estimation method	OLS	OLS	OLS
Adjusted R-squared	0.961	0.718	0.960

Note: The model-implied net supply measure and the yield-curve factors are estimated from the term structure model. ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. t-statistics are presented in the brackets.

sample period is limited to the three sub-periods under the ZLB regime, in column (iii). The estimated coefficient on the net supply measure is positive but insignificant at the 10 percent critical level. In sum, the net supply of relatively longer-term (but less than 10 years) JGBs tends to widen the 10-year/3-year spread, which is consistent with the preferred-habitat theory. However, the result of the above sub-sample regression as well as the calculated asymmetry in the model-implied net supply measure between the two regimes imply that the net supply effect was not stronger in the zero lower bound periods, in contrast to the regression analysis in Section III.[43]

B. Regression of the 20-year/10-year Yield Spread

Next, we examine the effects of the model-implied net supply measure on the 20-year/10-year spread. As before, we first check how the change in the maturity structure (z_{nt} for all n) is related to the net supply measure in the 20-year/10-year spread equation (5), using the data on JGB's maturity structure described in Section II for z_{nt}. If the arbitragers extend their maturity by buying 20-year bonds and selling 1-year bonds by a 1 percent share of their total JGB holdings (at the same time, the preferred-habitat investors shorten their maturity by selling 20-year bonds and buying 1-year bonds) under the normal regime, the net supply measure widens the 20-year/10-year yield spread

43. To address the possible endogeneity problem mentioned in Section III, we also conduct two-stage least square estimation using the same instrumental variables as in Section III or the lags of the model-implied net supply measure. The results are generally unchanged from those using the original (not instrumented) model-implied net supply measure shown in Table 11.

Table 12 Regression of 20-year/10-year Yield Spread from the Term Structure Model Approach

Dependent variable: 20-year/10-year JGB zero-coupon yield spread

	(i)	(ii)	(iii)
Model-implied net supply measure	1643.433 *** [5.40]	1582.461 *** [7.03]	1838.714 *** [2.66]
Yield-curve factor 1	0.006 [0.83]	0.007 * [1.78]	0.017 [1.34]
Yield-curve factor 2	–0.042 *** [–4.92]	–0.016 *** [–6.19]	–0.014 [–1.33]
Yield-curve factor 3	0.026 *** [4.53]	0.001 *** [4.21]	0.001 ** [2.10]
ZLB dummy × yield-curve factor 1	0.009 [0.86]		
ZLB dummy × yield-curve factor 2	0.027 *** [2.87]		
ZLB dummy × yield-curve factor 3	–0.025 *** [–4.38]		
Constant	0.314 *** [4.58]	0.380 *** [7.17]	0.358 ** [2.24]
ZLB dummy	0.087 *** [2.90]		
Sample period	1992M01-2014M09	1992M01-2014M09	1999M03-2000M07 2001M03-2006M06 2008M12-2014M09
Estimation method	OLS	OLS	OLS
Adjusted R-squared	0.343	0.287	0.439

Note: The model-implied net supply measure and the yield-curve factors are estimated from the term structure model. ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. t-statistics are presented in the brackets.

by about 0.44 basis point, given the estimated term structure model coefficients and assuming the parameter value of the arbitragers' risk aversion to be $\gamma = 100$. If, on the other hand, the arbitragers shorten their maturity (the preferred-habitat investors extend their maturity) by carrying out the opposite transaction under the ZLB regime, the net supply measure tightens the 20-year/10-year yield spread by about 0.22 basis point. Although the asymmetry between the two regimes is smaller than in the 10-year/3-year spread, the net supply effect on the 20-year/10-year spread is again stronger under the normal regime than under the ZLB regime, given the degree of the arbitragers' risk aversion.

Then, given the model-implied net supply measure and yield-curve factors, we estimate equation (5) by OLS. As summarized in Table 12, in both specifications with and without the coefficient dummies for the ZLB regime on the yield-curve factors, the estimated coefficient on the net supply measure, γ, is significantly positive. However, the size of the estimated γ, in both specifications, differs substantially from the results obtained from the 10-year/3-year spread equation. When the sample period is limited to the three sub-periods under the ZLB regime, the estimated γ is significantly positive and its size is slightly larger than those obtained from both the specifications in the full sample period. In sum, the net supply of relatively longer-term JGBs tends to widen the 20-year/10-year spread, which is again consistent with the preferred-habitat theory. However, the result of the above sub-sample regression combined with the calculated asymmetry in the model-implied net supply measure between the two regimes imply

Table 13 Regression of Holding-Period Excess Returns from the Term Structure Model Approach

	Dependent variable: ex-post 1-year holding excess return of 10-year JGBs			Dependent variable: ex-post 3-year holding excess return of 10-year JGBs		
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
Model-implied net supply measure	–9.158 [–1.14]	7.807 *** [4.22]	36.828 [0.34]	–4.697 *** [–2.89]	5.961 *** [13.54]	14.226 [0.94]
Yield-curve factor 1	0.355 * [1.79]	0.564 *** [5.41]	0.808 *** [2.92]	0.449 *** [11.24]	0.302 *** [12.06]	0.433 *** [10.02]
Yield-curve factor 2	1.552 *** [7.53]	0.290 *** [4.28]	0.301 ** [2.56]	0.408 *** [9.84]	0.170 *** [10.07]	0.274 *** [10.82]
Yield-curve factor 3	–0.853 *** [–6.23]	0.018 ** [2.32]	0.027 ** [2.39]	–0.218 *** [–7.91]	0.011 *** [5.76]	0.014 *** [8.20]
ZLB dummy × yield-curve factor 1	0.463 * [1.76]			–0.009 [–0.15]		
ZLB dummy × yield-curve factor 2	–1.242 *** [–5.46]			–0.130 *** [–2.68]		
ZLB dummy × yield-curve factor 3	0.878 *** [6.39]			0.230 *** [8.35]		
Constant	8.654 *** [2.76]	1.844 *** [3.22]	0.698 [0.21]	6.080 *** [9.64]	1.524 *** [10.98]	0.932 ** [2.03]
ZLB dummy	–6.589 ** [–2.24]			–4.588 *** [–7.73]		
Sample period	1992M01-2013M09	1992M01-2013M09	1999M03-2000M07 2001M03-2006M06 2008M12-2013M09	1992M01-2011M09	1992M01-2011M09	1999M03-2000M07 2001M03-2006M06 2008M12-2011M09
Estimation method	OLS	OLS	OLS	OLS	OLS	OLS
Adjusted R-squared	0.422	0.310	0.273	0.857	0.761	0.641

Note: The model-implied net supply measure and the yield-curve factors are estimated from the term structure model. ***, **, and * indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels, respectively. t-statistics are presented in the brackets.

that the net supply effect in the zero lower bound periods was not necessarily stronger than that in the full sample period.

C. Regression of Holding-Period Excess Returns

We then examine the effects of the model-implied net supply measure on the holding-period excess returns, the third term on the right-hand side of equation (6). As shown in the analysis of the one-period holding excess return in Koeda (2015), the net supply measure increases when the average maturity of JGBs held by the arbitragers is extended, but there is asymmetry in this relationship between the two regimes: the net supply measure is more sensitive to the arbitragers' average maturity under the normal regime than under the ZLB regime.

Keeping this relationship in mind, given the model-implied net supply measures and yield-curve factors, we estimate equation (6) using the actual data on the ex-post 1-year and 3-year holding excess returns of 10-year JGBs by OLS. The results are summarized in Table 13. Regarding the 1-year holding excess return, the estimated coefficient on the net supply measure, γ, is significantly positive only in the specifications without the coefficient dummies for the ZLB regime on the yield-curve factors. Even in this specification, the size of the estimated γ differs substantially from the results obtained from the term spread equations as well as the other specification of the holding-period excess returns. When the sample period is limited to the three sub-periods under the ZLB regime, the estimated γ is positive but insignificant. Results similar to these are also obtained for the 3-year holding excess return. Therefore, we could not obtain a robust and reliable estimate of γ to examine the net supply effects on the holding-period excess returns from this term structure model approach.[44]

V. Effects of the BOJ's JGB Purchases

Based on the analyses in Sections III and IV, we can extract the effects of the BOJ's purchases of JGBs and maturity extension as part of the QQE, which started in April 2013, on term spreads.[45] The idea is simple: we just plug in the contributions of the BOJ's JGB purchases and/or maturity extension to the various net supply measures in the regression equations in the preceding two sections, and calculate the cumulative net supply effects through our sample period after the implementation of QQE based on the estimated coefficients. We briefly summarize and compare these calculation results in the latter part of this section.

In general, however, it is difficult to identify the effects of a central bank's bond purchases on long-term interest rates in the data. As mentioned in Introduction, there are various channels of transmission of the effects including those that are not captured by our analyses based on the preferred-habitat theory. Moreover, the timing and persistence of the effects are hard to identify. As shown in Figure 6 in Section II, long-term interest rates actually started to decline before the QQE was implemented. This was probably because market participants had widely anticipated since the Prime Minister Abe took office near the end of December 2012 that he would appoint a new BOJ governor who would take a more aggressive monetary easing policy, including the BOJ's purchases of more and longer-term JGBs, compared with the governor at the time whose term was going to expire in April 2013.[46] These "anticipated" net supply effects are not captured, at least directly, in our analyses. Regarding the persistence of the net supply effects, it should be distinguished between the instantaneous response of bond prices to a central bank's ongoing purchase operations (the "flow effects") and persistent changes in bond prices resulting from movements along the bond demand curve (the "stock effects"), as investigated by D'Amico and King (2013). While our analyses using monthly data in the preceding sections focus mainly on the stock effects, the flow effects may also be important, as they could reflect some impairments in liquidity and functioning of the government bond markets as well as updates of market participants' expectations about the details of the central bank's operations.

In what follows, before showing the results calculated from our analyses in the preceding sections, we conduct a simple event-study analysis of the BOJ's outright purchases of JGBs using daily data, as a reference for comparison. While many event-study analyses in the literature focus on a number of official announcements regarding the framework of asset purchase programs that affect market participants' expectations about the net supply of the government bonds, we cannot do so for this analysis of the QQE because there was no major change in the framework of JGB purchase program in the QQE after the initial announcement on April 4, 2013 until the end of our sample period. Instead, we include each offer date of market operations by the BOJ in our

44. Due to this difficulty in estimating γ, Koeda (2015) reports her results with some alternative values of γ.
45. We do not calculate the effects of the BOJ's JGB purchases on holding-period excess returns because the data near the end of our sample period (1 year or 3 years) are not available.
46. The anticipations of market participants in this period are discussed in, for instance, Ito (2014).

event set. Therefore, our event-study analysis below may mainly capture the flow effects rather than the announcement effects of the BOJ's JGB purchases.

A. Event-Study Analysis

As mentioned above, our baseline event set includes all dates of offer of the BOJ's outright purchases of JGBs during the period between April 4, 2013 and September 30, 2014, which amounts to 173 events.[47] We also use an alternative event set that focuses on the outright purchases of JGBs with maturities over 10 years, which amounts to 87 events. As long-term interest rates started to decline before the QQE was implemented as mentioned above, we also examine an extended period from January 4, 2013 to September 30, 2014, in which the baseline events and alternative events increase to 185 and 90, respectively.[48]

Regarding the response window, we use one-day and two-day windows around the offer date: the former is measured from the closing level of the working day prior to the offer date to the closing level of the offer date, and the latter is measured from the closing level of the two working days prior to the offer date to the closing level of the offer date. We use the two-day window because many market participants anticipate the details (purchase amount, JGB maturity, etc.) of the operation to some extent in the working day prior to the offer date. At the same time, however, setting a wider window increases the risk of including effects on long-term interest rates that are unrelated to the BOJ's purchases. We calculate the cumulative changes in long-term interest rates during the response window of each event as a measure of the effects of the BOJ's JGB purchases. Specifically, we examine the effects on the 10-year yield, 20-year yield, 10-year/3-year spread, and 20-year/10-year spread on JGBs, which are all the zero-coupon yields as in the analyses in the preceding sections.

The results are summarized in Table 14. The upper table shows the results for the extended period from January 2013, and the lower table shows the results for the period after the implementation of QQE in April 2013. As there is only a small number of events in the period between January and April 2013 (12 in the baseline event set and 3 in the alternative event set), the cumulative changes during the response windows are generally not substantially different between the upper and lower tables. In contrast, the total cumulative changes through the two periods are very different, as shown in the bottom line in each table, because the zero-coupon yields decreased substantially in the period between January and April 2013 (23.8 basis points for 10-year yield) and then lowered only slightly (3.1 basis points) after the implementation of QQE until the end of our sample period.[49] Therefore, while the effects (cumulative changes during the response windows) in the extended period (shown in the upper table) are generally comparable with the total cumulative changes in the same period, the effects in the period after the implementation of QQE (the lower table) are much larger than the total cumulative changes in the corresponding period. The latter implies that the

47. The corresponding dates of exercise of the operations are generally two working days after the offer dates.
48. The market operations conducted in the period prior to April 4, 2013 include those under the asset purchase program in "Comprehensive Monetary Easing" as well as those conducted as a regular operation tool.
49. After the end of the sample period (September 30, 2014), the long-term interest rates decreased substantially again, especially after the expansion of QQE was announced on October 31, 2014. The decrease in the 10-year zero-coupon yield from September 30, 2014 to March 31, 2015 was 13.2 basis points.

Table 14 Event-Study Analysis

(1) From January 4, 2013 to September 30, 2014 (Extended period)

bps

Events	Windows	10-year yield	20-year yield	10-year/3-year spread	20-year/10-year spread
Offer of the outright purchases of JGBs operations by BOJ (185 events)	Two-day	−46.5	−57.6	−35.8	−11.1
	One-day	−27.4	−30.6	−16.4	−3.2
Offer of the outright purchases of JGBs operations with maturities over 10 years by BOJ (90 events)	Two-day	−56.5	−73.2	−52.5	−16.7
	One-day	−41.8	−46.3	−32.3	−4.5
Total changes through the period (427 days)		−26.9	−43.8	−25.6	−16.9

(2) From April 4, 2013 to September 30, 2014

bps

Events	Windows	10-year yield	20-year yield	10-year/3-year spread	20-year/10-year spread
Offer of the outright purchases of JGBs operations by BOJ (173 events)	Two-day	−32.8	−35.6	−23.8	−2.8
	One-day	−23.2	−23.0	−11.1	+0.2
Offer of the outright purchases of JGBs operations with maturities over 10 years by BOJ (87 events)	Two-day	−49.9	−71.4	−46.2	−21.5
	One-day	−40.7	−48.6	−30.8	−7.9
Total changes through the period (367 days)		−3.1	−3.0	−4.7	+0.1

Note: The table shows the cumulative changes in the zero-coupon yields and spreads during the response window of each events. One-day window is measured from the closing level of the working day prior to the offer date to the closing level of the offer date, and two-day window is measured from the closing level of the two working days prior to the offer date to the closing level of the offer date.

interest rates actually increased on non-event days reflecting many factors including rising inflation expectations, a rebound in the economic outlook, and improvements in market risk sentiment.[50]

From the comparison between the alternative event sets and response windows, we can see that the effects in the two-day window of the alternative (narrower) event set are the largest. This implies that the interest rates decreased remarkably on the day before the offer date of the outright purchases of JGBs with maturities over 10 years. From the comparison between the interest rate measures, it is shown that the effects on the 20-year yield are larger than those on the 10-year yields, while the effects on the 10-year/3-year spread are larger than those on the 20-year/10-year spread. These relationships are generally consistent with the total cumulative changes through the extended period from January 2013.

B. Results from Regression and Term Structure Model Approaches

We then show the results calculated from the analyses in Sections III and IV for the effects of the BOJ's JGB purchases and maturity extension on term spreads. As mentioned above, we plug in the BOJ's contribution to the various net supply measures in the regression equations and calculate the monthly cumulative net supply effects. The

50. Similar results of the decomposition between event days and non-event days are also obtained in event-study analyses using the U.S. data, such as Gagnon *et al.* (2011).

Table 15 Effects of the BOJ's JGB Purchases

(1) From January 2013 to September 2014 (Extended period)

bps

	Regression approach				Model approach	Event-study analysis	Actual changes through the period
	Baseline	Alternative specification	Sub-sample regressions				
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
JGB term spread (10-year/3-year)	−11.0	−22.9	−16.5	−41.9	−9.7 to −2.1	−52.5 to −16.4	−25.6
JGB term spread (20-year/10-year)	−27.7	−16.8	−25.0	−25.2	−8.5 to −8.1	−16.7 to −3.2	−16.9
Term premium (10-year)	−38.0 <−60.0>	—	—	—	—	—	−27.0

(2) From April 2013 to September 2014

bps

	Regression approach				Model approach	Event-study analysis	Actual changes through the period
	Baseline	Alternative specification	Sub-sample regressions				
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
JGB term spread (10-year/3-year)	−10.9	−22.6	−16.4	−41.5	−9.2 to −2.0	−46.2 to −11.1	−4.7
JGB term spread (20-year/10-year)	−27.8	−16.9	−25.0	−25.3	−7.9 to −7.6	−21.5 to +0.2	+0.1
Term premium (10-year)	−37.6 <−56.1>	—	—	—	—	—	−2.0

Note: In the regression approach, the effects of the BOJ's maturity extension are presented, except that the values in the angular brackets are the effects of the increase in the JGB purchases as well as the maturity extension.

results are summarized in Table 15. The upper table shows the results for the extended period from January 2013, and the lower table shows the results for the period from April 2013. As in Table 14, the calculated cumulative effects are generally not substantially different between the upper and lower tables because the BOJ's holding of JGBs and its average maturity changed little in the period between January and April 2013.

The results for the term spreads from the regression approach in Section III are calculated using the estimated coefficients in the baseline and an alternative specification (the baseline estimation shown in Tables 2 and 6 and the alternative specification using the BOJ's average maturity separately as shown in columns (iv) of Tables 3 and 7), and two sub-sample regressions with the baseline specification (the zero lower bound periods and the period after the global financial crisis as shown in columns (viii) and (ix) of Tables 5 and 9). In the baseline specification, we calculate the pure contribution of the BOJ's maturity extension (not including the contribution of changes in its share of the outstanding JGBs) to the arbitragers' average maturity, which is represented by the white bars in Figure 5 in Section II, and then multiply it by the estimated coefficients on the (instrumented) arbitragers' average maturity.

Regarding the effect on the 10-year/3-year spread, the result using the baseline estimation is 11.0 basis points for the extended period, as shown in column (i) of Table 15. The result from the alternative specification (column (ii)) is larger, possibly reflecting the downward bias caused by the endogeneity problem remaining in the baseline result. The results from the sub-sample regressions (columns (iii) and (iv)) are also larger than the baseline result, reflecting the larger estimated coefficients. Regarding the effect on the 20-year/10-year spread, the result using the baseline estimation is 27.7 basis points for the extended period. In contrast to the effects on the 10-year/3-year spread, however, the result from the alternative specification is smaller than the baseline result. As a result, within these alternative specifications, the effect on the 10-year/3-year spread is larger than that on the 20-year/10-year spread, while the former is smaller using the baseline estimation. Overall, these results obtained from the regression approach are broadly comparable with the range resulting from the event-study analysis (column (vi)) and the actual changes in the term spreads through the extended period from January 2013 (column (vii)).

The results for the term spreads from the term structure model approach in Section IV are calculated using the estimated coefficients in the specification with the coefficient dummies for the ZLB regime on the yield-curve factors (columns (i) of Tables 11 and 12). We calculate the contribution of changes in the maturity structure of JGBs held by the BOJ to the model-implied net supply measure, and then multiply it by the estimated coefficients on the net supply measure. In addition, we also take account of the effects of changes in the yield-curve factors on the term spreads under certain assumptions about the relationship between the maturity structure and the yield-curve factors.[51] The results are shown in column (v) of Table 15 as ranges depending on the assumption of the above relationship. In both the results for the 10-year/3-year and 20-year/10-year term spreads, the ranges of the effects are relatively small compared with the results from the regression approach and the event-study analysis. This may reflect the model estimation results in which the yield-curve factors are less volatile and the yields are less responsive to the yield-curve factors under the ZLB regime.

Finally, we calculate the effect of the BOJ's JGB purchases on the 10-year term premium derived from the term structure model in Section IV. The term premium is defined as the actual bond yield minus the average expected future short-term interest rates over the life of the bond.[52] Then the derived term premium is regressed on the arbitragers' average maturity, the change in the ratio of the arbitragers' holding JGBs (less than 10 years) to nominal GDP, the equity premium, the consumer price index inflation, the output gap, and the constant.[53] The estimated coefficients on the two net supply measures, the arbitragers' average maturity and the change in the arbitragers' debt-to-GDP ratio, are both significantly positive. Using these estimated coefficients, we calculate the effects of the BOJ's JGB purchases and maturity extension on the 10-year term premium. The result is shown in the bottom line in Table 15. The effect of the maturity extension is 38.0 basis points and the effect of the increase in the JGB purchases as well as the maturity extension is 60.0 basis points (both for the extended

51. Specifically, we assume that the correlation between the maturity structure that was most influenced by the BOJ's JGB purchases and the first yield-curve factor is in the range between 0.1 and 0.6, judging from the data.
52. To calculate the expectations components in the n-period term premium, we obtain $1, 2, \ldots, n$ period forecasts of the future short-term interest rates at each month via two-regime three-variable VAR forecasting. See Appendix 2 for details.
53. The estimation result is as follows:
$$tp_t^{10} = 0.764\,[3.20]x_t^{10} + 118.2\,[3.23]xxx_t^{10} - 0.287\,[-3.31]z_t^3 + 0.076\,[0.76]z_t^5 + 0.067\,[1.11]z_t^6 - 1.660\,[-1.49] + u_t,$$
where tp_t^{10} is the 10-year term premium, and x_t^{10} and z_t^3 are the same as in equation (1). xxx_t^{10} is the change in the ratio of JGBs (less than 10 years) held by the arbitragers to nominal GDP (smoothed by the Hodrick–Prescott filter), z_t^5 is the consumer price index inflation, and z_t^6 is the output gap estimated by the OECD. The values in the brackets are the Newey–West adjusted t-statistics. The adjusted R-squared is 0.497. The sample period is from January 1992 to September 2014.

period). These results are broadly comparable with the cumulative changes in the 10-year yield calculated in the event-study analysis shown in Table 14 (27.4 to 56.5 basis points for the extended period).[54]

VI. Concluding Remarks

In this paper, we have examined how the net supply of JGBs affected the term structure of interest rates and the risk premium on long-term JGBs using our constructed database on the amount outstanding of JGBs categorized by holder and remaining maturity. Both approaches using single-equation regressions and a term structure model confirm that the net supply of JGBs had significant effects on long-term interest rates, which is consistent with the preferred-habitat theory. The regression approach implies that the net supply effects were stronger in the zero interest rate periods, while this relationship was not found using the model approach. We also calculated the net supply effects of the BOJ's JGB purchases as part of the QQE using both approaches and compared the results with those obtained from a simple event-study analysis.

Based on our analyses in this paper, several directions for future research could be pursued. First, we have left the problem of when and how the effectiveness of the net supply of JGBs on long-term interest rates changed. A clear solution to this problem would provide useful implications for the future normalization of unconventional monetary policy as well as the government debt management policy. Second, there are several channels through which a shock to the net supply of JGBs affects long-term interest rates. It would be essential to empirically disentangle these channels for more accurately measuring the size, timing, and persistence of the net supply effects. On the theoretical side, there would be more room for improvement in specifications of the preferred-habitat theory, including a consideration of some heterogeneity among preferred-habitat investors, and for developing other theories on the effects of a central bank's purchases of government bonds on long-term interest rates.[55] Third, while we focus mainly on the gradual and persistent changes in interest rates using monthly data, daily or intra-day fluctuations and instantaneous responses to a net supply shock such as the effect of a central bank's ongoing purchase operations are also important, as discussed in Section V. In this regard, how the scarcity channel captured in our analyses relates to short-term liquidity would be an important problem left for future consideration. Finally, given the significant net supply effects on long-term interest rates, some normative and policy implications could be derived from this result. The current situation in Japan is that the BOJ extends the maturity of its holding of JGBs to achieve its price stability target and the private preferred-habitat investors extend their maturity to match their long-duration liabilities, while the government extends the maturity of its

54. BOJ (2015) estimates the effects of the BOJ's JGB purchases on real interest rates using several approaches. According to its regression approach, the cumulative effect from March 2013 to December 2014 was minus 80 basis points in terms of 10-year yields. This result seems broadly comparable with our results on the nominal 10-year term premium, considering that the long-term inflation expectations of economists and market participants increased by around 40 to 50 basis points during the corresponding period.

55. For instance, Williamson (2014) constructs a model in which short-maturity government debt has a greater degree of pledgeability than long-maturity government debt and collateral is collectively scarce, in which case a central bank's purchases of long-maturity government debt could flatten the yield curve.

debt to reduce the risks associated with refinancing. In terms of overall social welfare, however, there might be some room to improve the situation by pursuing better coordination among these entities.[56] We hope that our analyses in this paper will serve as a useful starting point for future research and policy discussions.

56. Greenwood *et al*. (2014) discuss the government debt management policy in the U.S. and suggest revised institutional arrangements to promote greater cooperation between the Treasury and the Federal Reserve when the conventional monetary policy is constrained by the zero lower bound.

APPENDIX 1: Database on JGBs

In this appendix, we explain how to construct the database on the amount outstanding of JGBs categorized by holder and remaining maturity.

A type of JGBs that is used in our analyses throughout the paper is fixed-rate JGBs, which cover the majority of the government debt in Japan (around 75 percent of the central government bonds and 60 percent of the general government gross debt as of 2013). In particular, we use 2-year, 4-year, 5-year, 6-year, 10-year, 20-year, 30-year, and 40-year fixed-rate JGBs. We do not use 15-year floating-rate bonds, inflation-linked bonds, financial bills, Treasury bills (T-bills), or any other discount bonds. We exclude directly underwritten bonds and focus on marketable bonds. Besides the total issuance of JGBs, we consider three sectors: the Bank of Japan (BOJ), insurance companies, and pension funds. Regarding the maturity structure, we use face-value outstanding of JGBs issued/held by each sector,[57] and categorize them into six categories along with their remaining maturities:[58] "less than or equal to 1 year," "longer than 1 year but less than 3 years," "longer than 3 years but less than 5 years," "longer than 5 years but less than 7 years," "longer than 7 years but less than 10 years," and "longer than 10 years." Our database has the form of a "sector (four)" by "maturity category (six)" matrix for every end of month from January 1992 to September 2014, which corresponds to the main sample period of our analyses. Due to unavailability of data, for some sectors and some sample periods we estimate the maturity structure using the data on the total amount outstanding of JGBs held by the sectors. The overview of our database and data availabilities is summarized in Table 16.

Using the above database, we calculate the average maturity of JGBs held by the arbitrager, which is used as a key net supply measure in our analyses. As explained in Section II, we define "arbitragers" as investors excluding the BOJ, insurance companies, and pension funds. The average maturity of JGBs held by arbitragers is calculated as the weighted sum (or difference) of the average maturity of the total issuance of JGBs, the BOJ's average maturity, insurance companies' average maturity, and pension funds' average maturity.[59] The weights are based on the share of JGBs held by each sector. As explained later, our maturity structure data only cover the period from September 2003 to September 2014 due to the data limitations. To calculate the arbitragers' average maturities between January 1992 and August 2003, we assume that the maturity structure of JGBs held by each sector is exactly the same as that for the total issuance. We calculate the arbitragers' average maturity using the step adjustment method for the period, and connect it with the data from September 2003 to September 2014.

In the following, we explain the detail of data construction method sector by sector.

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57. To avoid the endogeneity problem in our regression analyses, we basically use the data in face-value terms. However, we cannot obtain some data of this type. In this case, we estimate it based on the data in market-value terms. The details are explained later.
58. It is possible to calculate the remaining maturities more precisely for the total issuance and JGBs held by the BOJ. We use these six categories because data for insurance companies are only available in these categories.
59. The average maturity of JGBs with maturities less than 10 years held by the arbitragers are calculated in the same way, except that data with maturities less than 10 years are used.

Table 16 Overview of Our JGB Database and Data Availabilities

	Total issuance	Bank of Japan	Insurance companies				Pension funds
				Private life insurance companies	Nonlife insurance companies	Japan Post Insurance	
Total amount of JGBs (face-value terms)	Available	Available from May 2001. Before May 2001, data are estimated based on the data in market-value terms provided by the flow of funds statistics.	Available as a sum of those of private life insurance companies, nonlife insurance companies, and Japan Post Insurance	Estimated based on the data in market-value terms provided by the flow of funds statistics. Discontinuity is adjusted.	Estimated based on the data in market-value terms provided by the flow of funds statistics. Discontinuity is adjusted.	Estimated based on the data in market-value terms provided by the flow of funds statistics, and disclosures. Discontinuity is adjusted.	Estimated (1) based on the data in market-value terms provided by the flow of funds statistics from 1997 to 2014 and (2) with assumption of a constant share from 1992 to 1997.
Maturity information	Available	Available from May 2001.	Available as a weighted average of those of private life insurance companies, nonlife insurance companies, and Japan Post Insurance.	Estimated (from Sep. 2003 to Sep. 2014) based on disclosures.	Estimated (from Sep. 2003 to Sep. 2014) based on disclosures.	Estimated (from Sep. 2003 to Sep. 2014) based on disclosures.	Estimated (1) based on the information from the benchmark index from May 2006 to May 2014 and (2) with an assumption of a constant trend from Sep. 2003 to Apr. 2006 and June 2014 to Sep. 2014.
Data sources	Japan Securities Dealers Association, "*Ko-Shasai Binran* (Japanese Bond Handbook)."	Bank of Japan, "Japanese Government Bonds held by the Bank of Japan", "Flow of Funds Statistics."	—	Disclosures by each company; Bank of Japan, "Flow of Funds Statistics."	Disclosures by each company; Bank of Japan, "Flow of Funds Statistics."	Disclosures by Japan Post Insurance; Bank of Japan, "Flow of Funds Statistics."	Disclosures by the Nomura Domestic Bond Index Fund Nomura BPI DC; Bank of Japan, "Flow of Funds Statistics."

A. Total Supply

We collect the issuance data from the *Ko-Shasai Binran* (Japanese Bond Handbook)[60] on every single JGB issued between January 1979 and September 2014.[61] This database provides data on the characteristics of each bond, including bond type, series number, date of issue, date of redemption, amount issued in face-value terms, and direct underwriter (if any).[62] To focus only on marketable bonds, we exclude the data with the direct underwriters. We also collect retirement by purchase data[63] from the Ministry of Finance to calculate the remaining JGBs outstanding.[64] The data include the bond type, series number, date of purchase, and the purchase amount in face-value terms. Subtracting retirement by purchase data from the corresponding bond issuance data, we calculate the remaining maturity and the remaining amount in face-value terms for every single JGB for each end of month,[65] and categorize the amount into six maturity categories. Finally, we sum over all remaining bonds to build up the total issuance data.

B. Bank of Japan

The BOJ has released statistics titled "Japanese Government Bonds held by the Bank of Japan" every month since June 2001.[66] These statistics report the information on JGBs held by the BOJ at the end of the previous month, including bond type, series number, and amount of holding in face-value terms. These statistics do not include JGBs that are directly underwritten by the BOJ.[67] We combine these data and the bond characteristics data explained in Appendix 1.A and, in the same way as the total supply, obtain a data series on the amount outstanding of JGBs with their maturity structures from May 2001 to September 2014.

As we cannot know the maturity structure of JGBs held by the BOJ prior to May 2001, we instead construct the data on the total amount outstanding of JGBs held by the BOJ between January 1992 and April 2001. The data in these periods are based on the flow of funds statistics,[68] which report the amount outstanding of JGBs held by the BOJ in market-value terms.[69] Using these statistics, we first interpolate the data linearly, and calculate the year-on-year growth rate of JGBs held by the BOJ for each month. Then we backwardly estimate the total amount outstanding of JGBs held by

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60. The handbook is published semiannually by the Japan Securities Dealers Association (available only in Japanese).
61. We exclude bonds redeemed before 1992.
62. For example, the Financial Investment and Loan Program is one of the largest direct underwriters in our early sample period.
63. We use only the data of retirement by purchase from the market. The data are available on the website of the Ministry of Finance as part of its "Notification of the Ministry of Finance" (available only in Japanese).
64. As prematurity redemption of fixed-rate JGBs has never been observed in Japan, the amount outstanding of marketable fixed-rate JGBs only decreases when the Ministry of Finance conducts a retirement by purchase other than the redemption at maturity.
65. We do not break the stream of each bond's cash flows into principal and coupon payments to construct the future cash flows as in Greenwood and Vayanos (2014).
66. The statistics are available on the website of the BOJ.
67. Before 1998, the BOJ directly underwrote JGBs to roll them over.
68. The statistics are available on the website of the BOJ.
69. The statistics include 15-year floating-rate bonds, inflation-linked bonds, and JGBs that were directly underwritten by the BOJ.

the BOJ in face-value terms given this growth rate and the data after May 2001.

C. Insurance Companies

The data on the total amount outstanding of JGBs held by insurance companies are based on the flow of funds statistics. Since the data in flow of funds statistics are in market-value terms, we avoid using them directly. The detail estimation method is as follows. After December 1997, we calculate the share of JGBs held by insurance companies in the net issuance of JGBs,[70] and multiply this by the total amount of issuance in the face-value terms we explained in Appendix 1.A. Before November 1997, due to the discontinuity in the statistics, we first calculate the year-on-year growth rate of JGBs held by insurance companies that are defined differently, and backwardly estimate the total amount outstanding of JGBs held by insurance companies given this growth rate and the data after December 1997.

As for the maturity structure, we estimate it based on the disclosures of individual insurance companies. Insurance business law in Japan requires every insurance company to disclose the amount outstanding of JGBs and T-bills[71] by period remaining at least once each business year.[72] In the data, companies report six categories of the amount outstanding of JGBs and T-bills based on the remaining maturity mentioned above. To estimate the maturity as precisely as possible, we consider separately the structure of private life insurance companies, nonlife insurance companies, and Japan Post Insurance.[73] As for the maturity structure of JGBs held by private life insurance companies, we sum up the data for 11 major life insurance companies, and obtain the shares of each maturity category in the total amount outstanding of JGBs and T-bills from September 2003 to September 2014, using linear interpolation and a seasonality adjustment method. The same method is applied for four major nonlife insurance companies and Japan Post Insurance, respectively.[74] We then multiply this by the corresponding total amount outstanding of JGBs of these three insurance categories, respectively,[75] and obtain the data on the amount outstanding of JGBs and T-bills categorized

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70. The net issuance of JGBs is defined as the total issuance of JGBs minus the amount of JGBs held by government agencies.
71. The evaluation method of JGBs and T-bills is mixed, and it differs along with the purpose of holding; bonds declared as held-to-maturity debt are evaluated in terms of face value, and other securities are evaluated in terms of market value.
72. Insurance companies disclose these data through annual or semiannual reports. Their reports are mostly available on their homepage, but many companies release them only in Japanese.
73. We treat Japan Post Insurance, a subsidiary of the Japan Post, as a preferred-habitat investor in this paper. As for Japan Post Bank, another subsidiary of Japan Post, we treat it as an arbitrager. Although its investment strategy may be different from that of other major banks, its average maturity is not substantially different from that of major domestic banks. In order to obtain marketable JGBs held by Japan Post Insurance, we subtract the amount outstanding of directly underwritten JGBs from the total amount outstanding of JGBs held by Japan Post Insurance. The amount outstanding of directly underwritten JGBs held by Japan Post Insurance is estimated based on the *Ko-Shasai Binran* (Japanese Bond Handbook).
74. Since it is almost impossible to collect data from every insurance company in Japan, we only use the data of major insurance companies, and assume them to be a representative value. We use the data of 11 major life insurance companies and four major nonlife insurance companies besides that of Japan Post Insurance.
75. The total amount outstanding of JGBs in each insurance category is calculated exactly the same way as we explained above, and the sum of the total amount outstanding of JGBs of these three categories is equal to the total amount outstanding of JGBs held by insurance companies. To obtain the total amount outstanding of JGBs held by Japan Post Insurance, we use its disclosure as well as the flow of funds, because of discontinuity

by maturity structure from September 2003 to September 2014. Finally, we sum up these three insurance categories, and subtract the total amount outstanding of T-bills held by insurance companies, which is based on the flow of funds statistics, from the amount outstanding of JGBs and T-bills with maturities of "less than or equal to 1 year" to obtain the amount outstanding of JGBs.

D. Pension Funds

The data on the total amount outstanding of JGBs held by pension funds are estimated based on the flow of funds statistics, which are in market-value terms as well. After December 1997, we calculate the share of JGBs held by pension funds plus public pensions in the net issuance of JGBs, and multiply this by the total amount we explained above. Before November 1997, due to the discontinuity in the statistics, we cannot obtain exact data on pension funds. We assume that the share between 1992 and 1997 is equal to the average share of JGBs held by pension funds in the total issuance between 1997 and 2014. Using this assumption, we obtain the data on the total amount outstanding of JGBs held by pension funds from January 1992 to September 2014.

As for the maturity structure, since there is no disclosure of JGBs' maturity structure information in pension funds, we use as a proxy the maturity structure information from the NOMURA Bond Performance Index, which is used by many pension funds as a benchmark index for passive management of JGBs.[76] The information is available through the disclosures of fund-of-funds style investment trusts in which the index is used as an underlying fund. The annual disclosure reports the information on JGBs held by the index, including bond type, serial number, and amount of holding in face-value terms.[77] We combine these data and the bond characteristics data explained in Appendix 1.A, and, in the same way as for total supply, obtain the data series on the amount outstanding of JGBs, categorized by maturity structure, using linear interpolation and a seasonality adjustment method, from May 2006 to May 2014. Since each category of maturity structure in these periods is slightly changed with a constant slope, we assume that the maturity structures from September 2003 to April 2006 as well as from June 2014 to September 2014 are equally changed like those between May 2006 and May 2014. Combining this assumption with the total amount outstanding of JGBs held by pension funds, we can also calculate the data on the amount outstanding of JGBs categorized by maturity structure in these periods.

in the statistics.

76. For example, the Government Pension Investment Fund declares that it uses this index as a benchmark in its annual disclosure.
77. We use information from the "Nomura Domestic Bond Index Fund Nomura BPI DC." The data are available through the Electronic Disclosure for Investors' NETwork (EDINET) system in Japan.

APPENDIX 2: Term Structure Model in Section IV

In relation to Section IV in the text, this appendix summarizes the underlying model and reports the estimated parameters in Table 17. For a full description of the model, derivations, estimation strategy, and discussions on estimated results, see Koeda (2015). Following Vayanos and Vila (2009), we assume that there are two types of agents in the government bond market: preferred-habitat investors and arbitragers. Arbitragers maximize the mean-variance expected returns on their government bond portfolio, while preferred-habitat investors prefer to hold particular maturities of government bonds. The bond market equilibrium price is determined by equating arbitragers' net demand and preferred-habitat investors' net supply for different maturities of bonds.

Specifically, the model follows a discrete-time version of Vayanos and Vila (2009) discussed by Hamilton and Wu (2012a, HW henceforth). It extends HW by (i) allowing the coefficients in yield-curve factor dynamics as well as the prices of risk to change at the zero lower bound (ZLB) regime to allow greater model flexibility, (ii) using latent factors instead of observable factors to improve model fit to the data, (iii) writing out the arbitragers' portfolio optimization problem at the ZLB regime solving for the bond market equilibrium price, and (iv) carrying out estimation using JGB data.

Bond pricing in "normal" time (in the absence of the ZLB) follows the same specification as HW. HW define the arbitragers' rate of return from period t to period $t+1$ on their portfolio ($r_{t,t+1}$) as:

$$r_{t,t+1} = \sum_{n=1}^{N} z_{nt} r_{n,t,t+1}, \tag{7}$$

where z_{nt} is a fraction of their portfolio in the bond of maturity n and $r_{n,t,t+1}$ is the holding-period return from period t to $t+1$ on the n-period bond.

At the ZLB, arbitragers face two types of regimes: the ZLB regime where the ZLB binds, and the normal regime where the ZLB does not bind (denoted by $s = 0$ and $s = 1$, respectively). They maximize the mean-variance expected returns weighted by the transition probability that the ZLB will continue to bind in the next period (π_{00}) or that it will be lifted (π_{10}). These transition probabilities (π_{i0} for $i = 0, 1$) are assumed to be exogenous and constant and add up to 1 ($\sum_{i=0,1} \pi_{i0} = 1$). The arbitragers' optimization problem at the ZLB regime can be expressed as:

$$\begin{aligned} \max_{z_{1t},\ldots,z_{Nt}} \sum_{i=0,1} \pi_{i0}[&E(r_{t,t+1}|s_{t+1} = i, s_t = 0, \mathbf{f}_t) \\ &- (\gamma/2)\, Var(r_{t,t+1}|s_{t+1} = i, s_t = 0, \mathbf{f}_t)], \end{aligned} \tag{8}$$

subject to

$$\sum_{n=1}^{N} z_{nt} = 1, \tag{9}$$

where γ captures the arbitragers' risk aversion. $\mathbf{f}_t$ are the yield-curve factors assumed to follow a VAR (1) process as:

$$\mathbf{f}_{t+1} = \mathbf{c}^i + \mathbf{\Phi}^i \mathbf{f}_t + \mathbf{\Sigma}^i \mathbf{u}_{t+1},$$

with normalization that gives the identity matrix $\mathbf{\Sigma}^1$ and a 3×1 vector of zeros $\mathbf{c}^1$ and, for parsimonious purpose, we assume that $\mathbf{\Sigma}^0$ is a diagonal matrix. We can approximately solve for $\mathbf{z}_t = (z_{2t}, \ldots, z_{Nt})'$, that is, the arbitragers' demand equation. $\mathbf{z}_t$ depends on expected excess one-period holding returns on different maturities of bonds. Given $\mathbf{z}_t$, the arbitragers' demand for the short-term bond, z_{1t}, can be derived by equation (9).

The preferred habitat's net supply of bonds (x_{nt}) is modeled in the same manner as the normal-time model, except that the supply-equation coefficients are allowed to take different values at the ZLB regime:

$$x_{nt} \equiv \varsigma^0_{t,n} - \alpha^0_n y_{t,n}, \text{ for } n = 1, \ldots, N, \tag{10}$$

where $y_{t,n}$ is the log yield and $\varsigma^0_{t,n} = \varsigma^0_n + \vartheta^0_n \mathbf{f}_t$.

By equating the demand and supply functions for each maturity of bonds, the equilibrium log bond prices can be derived as follows:

$$p^0_{t,n} = \bar{a}^0_n + \bar{\mathbf{b}}^0_n \mathbf{f}_t, \quad \text{for } n = 1, \ldots, N, \tag{11}$$

$$\bar{a}^0_n = \bar{a}^0_1 + \sum_{i=0,1} \pi_{i0} \left[\bar{a}^i_{n-1} + \bar{\mathbf{b}}^i_{n-1} \left\{ \mathbf{c}^i - \Sigma^i \lambda^i \right\} + (1/2)\, \bar{\mathbf{b}}^i_{n-1} \mathbf{\Sigma}^i \mathbf{\Sigma}^{i\prime} \bar{\mathbf{b}}^{i\prime}_{n-1} \right], \tag{12}$$

$$\bar{\mathbf{b}}^0_n = \sum_{i=0,1} \pi_{i0} \bar{\mathbf{b}}^i_{n-1} \left\{ \mathbf{\Phi}^i - \mathbf{\Sigma}^i \mathbf{\Lambda}^i \right\} + \bar{\mathbf{b}}^0_1, \tag{13}$$

with the prices of risk coefficients given by

$$\lambda^i = \gamma \mathbf{\Sigma}^{i\prime} \sum_{n=2}^{N} \bar{\mathbf{b}}^{i\prime}_{n-1} \left(\varsigma^0_n + \left(\alpha^0_n / n \right) \bar{a}^0_n \right), \ \mathbf{\Lambda}^i = \gamma \mathbf{\Sigma}^{i\prime} \sum_{n=2}^{N} \bar{\mathbf{b}}^{i\prime}_{n-1} \left(\vartheta^0_n + \left(\alpha^0_n / n \right) \bar{\mathbf{b}}^0_n \right).$$

Thus, at the equilibrium, bond prices in the ZLB regime follow the standard affine term structure model with regime shifts. Furthermore, by arbitragers' FOCs, it can be shown, the prices of risk can be expressed as a function of z_t, as follows:

$$\lambda^i_t \equiv \lambda^i + \mathbf{\Lambda}^i \mathbf{f}_t = \gamma \mathbf{\Sigma}^{i\prime} \sum_{n=2}^{N} z_{nt} \bar{\mathbf{b}}^{i\prime}_{n-1}. \tag{14}$$

Table 17 Estimated Parameters of the Term Structure Model in Section IV

Yield curve coefficients						Transition probability
$\bar{a}_1^1 \times 1200$	$\bar{\mathbf{b}}_1^1 \times 1200$			$\bar{a}_1^0 \times 1200$		π_{00}
0.02	0.02	0.07	0.18	0.00		0.988
(0.09)	(0.03)	(0.01)	(0.003)	(0.02)		(0.001)
Factor dynamics						
$\mathbf{c}^1$	$\mathbf{\Phi}^1$			$\mathbf{\Sigma}^1$		
0.00	0.85	0.09	–0.03	1.00	—	—
—	(0.82)	(0.06)	(0.32)	—	—	—
0.00	0.08	0.85	0.06	—	1.00	—
—	(0.05)	(0.82)	(0.17)	—	—	—
0.00	–0.05	0.13	0.87	—	—	1.00
—	(0.63)	(0.24)	(0.04)	—	—	—
$\mathbf{c}^0$	$\mathbf{\Phi}^0$			$\mathbf{\Sigma}^0$		
–0.11	0.93	0.16	0.00	0.65	—	—
(0.23)	(0.05)	(0.12)	(0.14)	(0.015)	—	—
–1.27	–0.11	0.76	0.01	—	0.64	—
(0.73)	(0.02)	(0.04)	(0.04)	—	(0.002)	—
10.01	0.49	1.56	0.87	—	—	11.98
(0.93)	(0.03)	(0.07)	(0.08)	—	—	(0.004)
$\mathbf{c}^Q$	$\mathbf{\Phi}^Q$			$\mathbf{\Sigma}_e^1 \times 1200$	$\mathbf{\Sigma}_e^0 \times 1200$	
–0.02	0.99	—	—	0.09	0.02	
(0.84)	(0.10)	—	—	(0.007)	(0.001)	
0.17	0.02	0.99	—			
(0.11)	(0.00)	(0.10)	—			
–0.15	–0.11	0.26	0.80			
(0.35)	(1.26)	(0.57)	(0.16)			

Note: $\mathbf{c}^1$ and $\mathbf{\Sigma}^1$ are normalized to be a 3×1 vector of zeros and a 3×3 identity matrix, respectively. Numbers in the parentheses are standard errors.

A. Risk Premium (Holding-Period Excess Returns)

At the bond market equilibrium, the model-implied expected k-period holding excess log return of j-period bonds under the normal regime ($rx^1_{t+1,j,k}$) can be expressed as:

$$E_t\left(rx^1_{t+1,j,k}\right) \equiv E_t\left(p^1_{t+1,j-k}\right) - p^1_{t,j} + p^1_{t,k}$$
$$= E_t\left(\bar{a}^1_{j-k} + \bar{\mathbf{b}}^1_{j-k}\mathbf{f}_{t+k}\right) - \bar{a}^1_j - \bar{\mathbf{b}}^1_j\mathbf{f}_t + \bar{a}^1_k + \bar{\mathbf{b}}^1_k\mathbf{f}_t,$$
$$= c^1_{j,k} + \mathbf{\Phi}^1_{j,k}\mathbf{f}_t + \gamma\bar{\mathbf{b}}^1_{j-1}\mathbf{\Sigma}^1\mathbf{\Sigma}^{1\prime}\sum_{n=2}^{N} z_{nt}\bar{\mathbf{b}}^{1\prime}_{n-1},$$

where $p^1_{t,j}$ is the j-period log price. The second equality holds because bond prices are modeled as an exponential affine function of the yield-curve factors. The third equality holds by using the recursive equations for $\bar{a}^1_j$ and $\bar{\mathbf{b}}^1_j$ and the arbitragers' FOCs, and also by solving for the conditional expectation term $\left(E_t\left(\bar{a}^1_{j-k} + \bar{\mathbf{b}}^1_{j-k}\mathbf{f}_{t+k}\right)\right)$ with $c^1_{j,k} \equiv \bar{a}^1_{j-k} + \bar{a}^1_k - \bar{a}^1_1 - \bar{a}^1_{j-1} - \frac{1}{2}\bar{\mathbf{b}}^1_{j-1}\mathbf{\Sigma}^1\mathbf{\Sigma}^{1\prime}\bar{\mathbf{b}}^{1\prime}_{j-1}$ and $\mathbf{\Phi}^1_{j,k} \equiv \left(\bar{\mathbf{b}}^1_k - \bar{\mathbf{b}}^1_1\right) + \bar{\mathbf{b}}^1_{j-k}\left(\mathbf{\Phi}^1\right)^k - \bar{\mathbf{b}}^1_{j-1}\mathbf{\Phi}^1$.

Similarly, the corresponding excess log return under the ZLB regime can be expressed as:

$$\begin{aligned}E_t\left(rx^0_{t+1,j,k}\right) &\equiv E_t\left(p^0_{t+1,j-k}\right) - p^0_{t,j} + p^0_{t,k}\\ &= E_t\left(\bar{a}^{s_{t+k}}_{j-k} + \bar{\mathbf{b}}^{s_{t+k}}_{j-k}\mathbf{f}_{t+k}\right) - \bar{a}^0_j - \bar{\mathbf{b}}^0_j\mathbf{f}_t + \bar{a}^0_k + \bar{\mathbf{b}}^0_k\mathbf{f}_t,\\ &= c^0_k + \mathbf{\Phi}^0_k\mathbf{f}_t + \gamma\sum_{i=0,1}\pi_{i0}\bar{\mathbf{b}}^i_{j-1}\mathbf{\Sigma}^{i\prime}\mathbf{\Sigma}^{i\prime}\sum_{n=2}^{N} z_{nt}\bar{\mathbf{b}}^{i\prime}_{n-1},\end{aligned}$$

where

$$\begin{aligned}c^0_k \equiv\ & \bar{a}^0_k - \bar{a}^0_1 + (\pi_{00})^k\left\{\bar{a}^0_{j-k} + \sum_{i=1}^{k}\left(\mathbf{\Phi}^0\right)^{i-1}\mathbf{c}^0\right\} + (1-\pi_{00})\,\bar{a}^1_{j-k}\\ &+ \sum_{l=1}^{k-1}(\pi_{00})^l(1-\pi_{00})\left(\bar{a}^1_{j-k} + \left(\mathbf{\Phi}^1\right)^{k-l}\sum_{i=1}^{l}\left(\mathbf{\Phi}^0\right)^{i-1}\mathbf{c}^0\right)\\ &- \sum_{i=0,1}\pi_{i0}\left[\bar{a}^i_{j-1} + \bar{\mathbf{b}}^i_{j-1}\mathbf{c}^i + (1/2)\,\bar{\mathbf{b}}^i_{j-1}\mathbf{\Sigma}^i\mathbf{\Sigma}^{i\prime}\bar{\mathbf{b}}^{i\prime}_{j-1}\right],\end{aligned}$$

$$\begin{aligned}\mathbf{\Phi}^0_k \equiv\ & \bar{\mathbf{b}}^0_k - \bar{\mathbf{b}}^0_1 + (\pi_{00})^k\,\bar{\mathbf{b}}^0_{j-k}\left(\mathbf{\Phi}^0\right)^k + (1-\pi_{00})\,\bar{\mathbf{b}}^1_{j-k}\left(\mathbf{\Phi}^1\right)^k\\ &+ \sum_{l=1}^{k-1}(\pi_{00})^l(1-\pi_{00})\,\bar{\mathbf{b}}^1_{j-k}\left(\mathbf{\Phi}^1\right)^{k-l}\left(\mathbf{\Phi}^0\right)^l - \sum_{i=0,1}\pi_{i0}\bar{\mathbf{b}}^i_{j-1}\mathbf{\Phi}^i.\end{aligned}$$

B. Term Spread

The model-implied j-period log bond price under the normal regime can be expressed as:

$$\begin{aligned}p^1_{t,j} &= c^1_{j-1} + \mathbf{\Gamma}^1_{j-1}\mathbf{f}_t - \bar{\mathbf{b}}^1_{j-1}\mathbf{\Sigma}^1\lambda^1_t, \quad \text{for } j = 1,\ldots,N,\\ &= c^1_{j-1} + \mathbf{\Gamma}^1_{j-1}\mathbf{f}_t - \gamma\bar{\mathbf{b}}^1_{j-1}\mathbf{\Sigma}^1\mathbf{\Sigma}^{1\prime}\sum_{n=2}^{N} z_{nt}\bar{\mathbf{b}}^{1\prime}_{n-1},\end{aligned}$$

where the first equality holds by the pricing equation (equations (11)–(13)) and the second equality holds by the arbitragers' FOCs with $c^1_{j-1} \equiv \bar{a}^1_{j-1} + \bar{\mathbf{b}}^1_{j-1}\mathbf{c}^1 + \frac{1}{2}\bar{\mathbf{b}}^1_{j-1}\mathbf{\Sigma}^1\mathbf{\Sigma}^{1\prime}\bar{\mathbf{b}}^{1\prime}_{j-1} + \bar{a}^1_1$ and $\mathbf{\Gamma}^1_{j-1} \equiv \bar{\mathbf{b}}^1_{j-1}\mathbf{\Phi}^1 + \bar{\mathbf{b}}^1_1$. This implies that the term spread between j- and k-period bonds ($j > k$) can be expressed as:

$$y_{t,j} - y_{t,k} = \tilde{c}^1_{j,k} + \tilde{\mathbf{\Gamma}}^1_{j,k}\mathbf{f}_t + \gamma\left((1/j)\,\bar{\mathbf{b}}^1_{j-1} - (1/k)\,\bar{\mathbf{b}}^1_{k-1}\right)\mathbf{\Sigma}^1\mathbf{\Sigma}^{1\prime}\sum_{n=2}^{N} z_{nt}\bar{\mathbf{b}}^{1\prime}_{n-1},$$

where $y_{t,j} = -p^1_{t,j}/j$, $\tilde{c}^1_{j,k} \equiv c^1_{k-1}/k - c^1_{j-1}/j$, and $\tilde{\mathbf{\Gamma}}^1_{j,k} \equiv (1/k)\,\mathbf{\Gamma}^1_{k-1} - (1/j)\,\mathbf{\Gamma}^1_{j-1}$.

Similarly, the j-period log bond price under the ZLB regime can be expressed as:

$$p^0_{t,j} = c^0_{j-1} + \mathbf{\Gamma}^0_{j-1}\mathbf{f}_t - \gamma\sum_{i=0,1}\pi_{i0}\bar{\mathbf{b}}^i_{j-1}\mathbf{\Sigma}^i\mathbf{\Sigma}^{i\prime}\sum_{n=2}^{N} z_{nt}\bar{\mathbf{b}}^{i\prime}_{n-1}, \quad \text{for } j = 1,\ldots,N,$$

where $c^0_{j-1} \equiv \bar{a}^0_1 + \sum_{i=0,1} \pi_{i0} \left(\bar{a}^i_{j-1} + \bar{\mathbf{b}}^i_{j-1} \mathbf{c}^i + (1/2)\, \bar{\mathbf{b}}^i_{j-1} \mathbf{\Sigma}^i \mathbf{\Sigma}^{i\prime} \bar{\mathbf{b}}^{i\prime}_{j-1} \right)$ and $\mathbf{\Gamma}^0_{j-1} \equiv \sum_{i=0,1} \pi_{i0} \bar{\mathbf{b}}^i_{j-1} \mathbf{\Phi}^i + \bar{\mathbf{b}}^0_1$. Then the term spread between j- and k-period bonds can be expressed as:

$$y_{t,j} - y_{t,k} = \tilde{c}^0_{j,k} + \tilde{\mathbf{\Gamma}}^0_{j,k} \mathbf{f}_t + \gamma \left(\sum_{i=0,1} \pi_{i0} \left(\bar{\mathbf{b}}^i_{j-1}/j - \bar{\mathbf{b}}^i_{k-1}/k \right) \mathbf{\Sigma}^i \mathbf{\Sigma}^{i\prime} \sum_{n=2}^{N} z_{nt} \bar{\mathbf{b}}^{i\prime}_{n-1} \right),$$

where $y_{t,j} = -p^0_{t,j}/j$, $\tilde{c}^0_{j,k} \equiv c^0_{k-1}/k - c^0_{j-1}/j$ and $\tilde{\mathbf{\Gamma}}^0_{j,k} \equiv (1/k)\, \mathbf{\Gamma}^0_{k-1} - (1/j)\, \mathbf{\Gamma}^0_{j-1}$.

C. Term Premium

We decompose the long-term bond yields into the expectations and term premium components. Following the typical definition in the literature, the term premium of an n-period bond yield is defined as the actual n-period bond yield minus the average expected future short-term interest rates over the life of the bond (i.e., $\left(\frac{1}{n}\right) E_t \left\{ \sum_{j=0}^{n-1} r_{1,t+j} \right\}$). To calculate the expectations components, we first obtain $1, 2, \ldots, n$ period forecasts of the future short-term interest rates at each month via two-regime three-variable VAR forecasting. Our model assumes that (i) if the initial regime is normal, then the corresponding VAR forecasting is reduced to a simple one-regime VAR forecasting. And (ii) if the initial regime is ZLB, then the regime shifts from ZLB to normal with the probability of $1 - \pi_{00}$ in the next period; once the regime shifts to normal, however, it remains normal for the rest of the forecasting period. This implies that there are $n + 1$ possible paths of regimes and expected future yield-curve factors when the initial regime is ZLB. The conditional expectation on n-period ahead yield-curve factors is computed by taking the average of possible values of expected yield-curve factors weighted by probability.

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