Mortgage convexity*

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Abstract

Most home mortgages in the United States are fixed-rate loans with an embedded prepayment option. When long-term rates decline, the effective duration of mortgage-backed securities (MBS) falls due to heightened refinancing expectations. I show that these changes in MBS duration function as large-scale shocks to the quantity of interest rate risk that must be borne by professional bond investors. I develop a simple model in which the risk tolerance of bond investors is limited in the short run, so these fluctuations in MBS duration generate significant variation in bond risk premia. Specifically, bond risk premia are high when aggregate MBS duration is high. The model offers an explanation for why long-term rates could appear to be excessively sensitive to movements in short rates and explains how changes in MBS duration act as a positive-feedback mechanism that amplifies interest rate volatility. I find strong support for these predictions in the time series of US government bond returns.

JEL classification: G10, G12.

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1. Introduction

A distinguishing feature of bond markets in the United States is the prominent role of fixed-rated home mortgages. Most fixed-rate mortgages issued in the US grant the borrower the right to prepay the loan at any time without penalty. When long-term interest rates decline, the option to prepay the mortgage and refinance at the current interest rate moves toward the money. Because lenders expect to be repaid sooner, the effective duration of outstanding mortgage-backed securities (MBS) falls; i.e., the sensitivity of MBS prices to changes in long-term yields declines. Conversely, as long-term interest rates rise, the prepayment option moves out-of-the-money and the effective duration of MBS rises. To quote a Wall Street adage, a mortgage-backed security "goes up like a two-year bond" when rates fall and "goes down like a six-year bond" when rates rise. More formally, the negative relationship between price and yield is convex for a noncallable bond. By contrast, the price-yield relation for callable bonds such as MBS is typically concave, a feature that bond investors call "negative convexity."

Aggregate mortgage refinancing activity varies significantly over time. Because refinancing entails large financial and nonfinancial costs, households typically follow a trigger rule, refinancing only when their option is sufficiently far in-the-money. Thus, depending on the past path of mortgage rates, there are times when many households move from being far from refinancing to being close to refinancing and vice versa. I argue that the resulting changes in aggregate MBS duration act as massive supply shocks to the total quantity of interest rate risk that must be borne by investors in the broader bond market. Because these shifts in bond market duration are large relative to bond investors' risk tolerance, they have a significant impact on equilibrium term premium; i.e., on the expected return on long-term default-free bonds over short-term default-free bonds. In summary, time variation in aggregate mortgage refinancing has a significant effect on the pricing of interest rate risk throughout the US fixed income market and, in particular, it has a large impact on US Treasury yields.

Why do shifts in expected household mortgage refinancing affect the aggregate amount of interest rate risk that bond market investors must bear? First, household borrowers only gradually exercise their prepayment options following a decline in prevailing mortgage rates. Second, household borrowers do not alter their bond holdings to hedge their time-varying interest rate exposure. That is, households do not adjust their asset portfolios to offset the time-varying interest rate risk they are bearing on the liability side. The gradual response of mortgage prepayments to changes in mortgage rates and the lack of household hedging means that households at times are bearing more or less interest rate risk. Conversely, bond investors at times are bearing less or more risk. Thus, shifting refinancing expectations generate a form of aggregate market congestion: sometimes most households are effectively borrowing long term, and other times they are borrowing short term. If the risk-bearing capacity of professional bond investors is limited in the short run, then the term premium must adjust to induce investors to bear these risks.

The sheer size of the MBS market within the US bond market plays a crucial role in this story. Bond markets have witnessed several MBS duration supply shocks that are larger than the shift in duration induced by the Federal Reserve's Quantitative Easing (QE) policies from 2008 to 2012.

For instance, on several occasions the quarterly change in MBS duration was equivalent to a $1 trillion increase (in 2012 dollars) in the supply of ten-year Treasury notes, with a corresponding reduction in the supply of short-term T-bills. By way of comparison, the quarterly new-issue supply of ten-year Treasury notes in 2012 was roughly $65 billion. Thus, past shifts in MBS duration have arguably been very large relative to the risk tolerance of bond market arbitrageurs.

I develop a simple model of this mortgage convexity mechanism. The model has the following key ingredients. First, expectations of future household mortgage refinancing vary over time, which induces shifts in the effective duration of household mortgage borrowing. Second, aggregate bond risks are priced by risk-averse, specialized bond investors, as in Vayanos and Vila (2009). This ensures that demand curves for aggregate bond risks slope downward, so the term premium must adjust to induce investors to absorb the aggregate supply of bond duration. Specifically, the model predicts that measures of aggregate bond market duration, which derive most of their power from variation in MBS duration, should positively predict bond excess returns. Furthermore, because shocks to MBS duration are fairly transient, one would expect fluctuations in MBS duration to lead to high frequency variation in bond risk premia.

The model shows how negative MBS convexity could contribute to the excess sensitivity of long-term yields to movements in short rates (Gürkaynak, Sack, and Swanson, 2005) and, more generally, to the excess volatility of long-term yields (Shiller, 1979). Excess sensitivity reflects what might be called an "MBS duration spiral." An initial shock to the short rate directly raises long yields due to the expectations hypothesis. The rise in long yields raises the duration of MBS. The term premium must rise to induce risk-averse bond investors to bear the larger aggregate quantity of interest rate risk. The resulting rise in yields further raises MBS duration, which further raises the term premium, and so on. Thus, the fact that MBS duration is increasing in long yields gives rise to a positive-feedback channel that generates excess sensitivity and excess volatility. The model also suggests that this positive-feedback mechanism is strongest when the mortgage market is on a refinancing cliff; i.e., when a small movement in long rates significantly impacts refinancing behavior and, hence, MBS duration. The model therefore predicts that excess sensitivity and volatility should be most pronounced when the MBS market is more negatively convex.

I also show that an MBS duration shock should have a larger effect on the expected excess returns of long-term bonds than on those of intermediate-term bonds. This is a natural consequence of the fact that an MBS duration shock raises the current duration risk premium in bond markets. However, because these shocks are transient, shocks to MBS duration have a humped-shaped effect on the yield curve and the forward rate curve. The effects of a supply shock on yields, which equals the effect on the bond's average returns over its lifetime, is greater for intermediate-term bonds than for long-term bonds when the supply shock is expected to be short-lived. That is, a shock to MBS duration increases the curvature of the yield curve. This suggests that transient shocks to MBS duration could account for some of the predictive power of the Cochrane and Piazzesi (2005) factor, which picks up time variation in the curvature of the yield curve and is useful for forecasting transitory variation in bond returns.

I find strong support for these predictions in US interest rate data between 1989 and 2012.

Measures of MBS duration are strong predictors of excess government bond returns. And they appear to contain information that is not reflected in traditional forecasting variables based on the current shape of the yield curve. My analysis of the time signature of these effects indicates that shocks to MBS duration have a transitory impact on term premia, which largely dissipates over the next 6 to 12 months. As predicted, shocks to MBS duration have a hump-shaped effect on the yield and forward rate curves. I also find that the excess sensitivity of long rates to short rates is more pronounced when the MBS market is on a refinancing cliff, i.e., when a move in long rates has a larger impact on aggregate MBS duration. Lastly, I find that option-implied volatility of long yields is higher when the MBS market is more negatively convex.

I lack a good instrument for MBS duration that would allow me to cleanly identify the demand curve for interest rate risk. As a result, I rely on predictive regressions that provide indirect evidence consistent with the view that shifts in MBS duration trace out a downward-sloping demand for interest rate risk. These return forecasting regressions are analogous to a regression of prices (term premia) on quantities (duration supply). Thus, I take considerable care to address several natural concerns raised by this indirect approach.

To begin, I control for an exhaustive set of factors that are thought to impact term premia to address concerns that my forecasting results are driven by an omitted variable. I then provide a host of further indirect evidence that is consistent with the MBS story, but that would not be predicted by alternative explanations for my findings. First, I show that the return forecasting power of MBS duration has grown significantly over time as MBS markets have grown relative to the rest of the US bond market. Second, I show that shifts in MBS duration have far more forecasting power for US bond returns than for foreign bond returns. Third, I show that shifts in MBS duration impact the spreads between corporate bonds, interest rate swaps, and Treasuries in precisely the way that one would expect if MBS investors were using Treasuries and swaps (but not corporate bonds) to hedge variation in MBS duration. Collectively, these additional findings provide further support for my story emphasizing shifts in MBS duration.

Bond investors often invoke shifts in MBS duration and the portfolio-hedging flows they trigger when explaining large movements in long-term interest rates. Many MBS investors delta-hedge the time-varying duration of MBS. If there is a drop in long rates that raises mortgage refinancing expectations and lowers MBS duration, these investors buy more long-term Treasuries, financed by selling some short-term T-bills, in order to maintain their prior asset duration. Given the prominence that these dynamics receive in practitioner commentary, their relative absence from the literature on the term structure of interest rates is somewhat surprising. The handful of exceptions includes Fernald, Keane, and Mosser (1994), Kambu and Mosser (2001), Perli and Sack (2003), Duarte (2008), and Malkhozov, Mueller, Vedolin, and Venter (2013). Kambu and Mosser (2001), Perli and Sack (2003), and Duarte (2008) each argue that these hedging flows impact interest rate volatility, whereas the primary focus of my paper is to investigate the impact of MBS duration on the equilibrium term structure of yields and term premia.[1] In

[1] Perli and Sack (2003) show that MBS hedging impacts swaption-implied yield volatility. Duarte (2008) also finds that MBS hedging forecasts high future realized volatility, which suggests that hedging flows not only impact interest rate option prices, but also impact the underlying bond yields.

contemporaneous work, Malkhozov, Mueller, Vedolin, and Venter (2013) also explore the impact of MBS duration on excess returns. Their study places more emphasis on the implications of MBS hedging for interest rate volatility, whereas I am primarily interested in understanding the pricing implications of these hedging flows.

The ideas in this paper connect to several broader strands of prior research. First, the idea that supply and demand effects can have important consequences in bond markets is central to a number of recent papers, including Vayanos and Vila (2009), Greenwood and Vayanos (2010, 2014), Krishnamurthy and Vissing-Jorgensen (2011, 2012), and Gagnon, Raskin, Remache, and Sack (2011). Important precursors to this recent work include Tobin (1958) and Modigliani and Sutch (1966). This literature has featured prominently in the portfolio balance effect interpretation of the Federal Reserve's Quantitative Easing policies, in which it is typically argued that Fed purchases of long-term assets reduce bond risk premia. Second, a vast literature is devoted to forecasting the excess returns on long-term bonds. Important contributions here include Fama and Bliss (1987), Campbell and Shiller (1991), and Cochrane and Piazzesi (2005). Third, a variety of papers find excess sensitivity, including Cochrane and Piazzesi (2002), Gürkaynak, Sack, and Swanson (2005), and Hanson and Stein (2012). Fourth, a number of papers argue that long-term yields are excessively volatile, including Shiller (1979), Perli and Sack (2003), and Duarte (2008). Finally, a growing literature, including Gabaix, Krishnamurthy, and Vigneron (2007), Greenwood and Vayanos (2014), and Vayanos and Vila (2009), explores the implications of limited arbitrage in fixed income markets.

The plan for the paper is as follows. Section 2 develops a simple model that generates several novel predictions. In Section 3, I find strong support for these predictions in the time series of US government bond returns. Section 4 concludes.

2. Theoretical framework

This section clarifies why shifts in MBS duration matter for bond pricing and develops a stylized model of the mortgage convexity mechanism.

2.1. Shocks to MBS duration

Suppose a large number of households announce that in six months time they will repay their existing mortgages at par and will take out new mortgages at the prevailing mortgage rate. And suppose for simplicity that this commitment to refinance in six months is binding. Suddenly, from the perspective of investors who own the existing mortgages, the mortgages behave just like short-term bonds, so mortgage holders are bearing less interest rate risk. And over the following six months, households are bearing significantly more interest rate risk. That is, they will be worse off if rates rise and better off if rates fall. To hedge this interest rate exposure, households could initially sell some long-term Treasuries with the intention of buying them back in six months.[2]

[2]The household would invest the proceeds in short-term bills for six months. If households were to hedge in this way, their desire to short long-term bonds would perfectly offset the temporary decline in mortgage duration.

However, due to a variety of frictions, costs, or a lack of financial sophistication, households do not hedge their time-varying exposure to interest rate risk. So, on net, a temporary reduction occurs in the amount of interest rate risk that fixed income investors must bear in equilibrium. And if the risk-bearing capacity of bond investors is limited, the expected excess return on long-term bonds must fall to equate the supply and demand for bonds.

In practice, households do not enter into binding agreements to refinance. Instead, household refinancing behavior is gradual, so shocks to long-term interest rates alter expected mortgage refinancing. Critically, I assume households do not hedge their interest rate exposure; i.e., households do not alter their bond holdings to hedge the time-varying interest rate risk they are assuming on the liability side of their balance sheets. This plausible assumption is the key friction that generates time-varying risk sharing between household borrowers and bond investors and, thus, shifts in aggregate bond market duration.

For instance, when future mortgage refinancings are expected to be high, households are effectively borrowing shorter term and, hence, are bearing more interest rate risk: households are exposed to the risk that they will refinance at a less advantageous rate. Conversely, when expected refinancings are low, households are effectively borrowing longer term, so investors are bearing more interest rate risk. The total quantity of interest rate risk borne in the economy remains unchanged. A single long-term asset, namely, housing, is going to be financed long term at some interest rate. What varies is the amount of interest rate risk that is borne by bond investors as opposed to household borrowers. Shifts in the extent of aggregate risk sharing between households and investors move the price of interest rate risk—the duration risk premium—in bond markets.

If the total quantity of interest rate risk in the economy is fixed, why do shifts in MBS duration move bond prices? One natural answer, as in Gabaix, Krishnamurthy, and Vigneron (2007), is that, while this could be a wash from the perspective of some representative household (e.g., when about to refinance, the representative household takes more interest rate risk on the liability side but is exposed to less interest rate risk on bond holdings), the change in the quantity of bond risk still looms large from the standpoint of a delegated investment manager who specializes in bond markets. If this manager is forced to invest a large fraction of her own wealth in the bond fund, e.g., to mitigate agency problems, then the manager perceives an increase in the total quantity of risk. Thus, risks that are idiosyncratic from an aggregate perspective could be priced because of the large undiversified exposures of specialized intermediaries who are the marginal buyers of bonds.[3]

Distinguishing between the interest rate risk of MBS and MBS model risk is crucial. Interest rate risk derives from the fact that the value of MBS depends on the level of rates, assuming that household prepayment behavior is a known deterministic function of rates. By contrast, model risk derives from the fact that household prepayment behavior is not a known function of interest rates. Instead, it is sometimes difficult to predict how prepayments will respond to

[3] A second answer is to think about an overlapping generations setting where households borrow when young and lend when old. Because the young are bearing less interest rate risk when refinancing expectations are low, this means that the old are bearing more risk, which impacts equilibrium term premia. In this telling, time-varying risk sharing between the old and young induces time variation in bond risk premia.

changes in rates. This paper focuses on the time-varying interest rate risk of the aggregate MBS market. The key insights are that household prepayment behavior is gradual and that households do not hedge their time-varying interest rate exposure. Thus, the total quantity of interest rate risk borne by professional bond investors can vary over time even if prepayment behavior is a deterministic function of the path of rates. As a result, the time-varying quantity of interest-rate risk emphasized here should be distinguished from MBS model risk as analyzed by Gabaix, Krishnamurthy, and Vigneron (2007).

2.2. Stylized model with short- and long-term bonds

The main predictions can be illustrated using a simple dynamic model that features only short- and long-term bonds at each date t. The expected excess return on long-term bonds over short-term bonds is the only endogenous variable at each date. The Internet Appendix extends the model to allow for multiple bond maturities.

A set of risk-averse arbitrageurs (e.g., fixed income hedge funds, fixed income dealers, etc.) price interest rate risk in bond markets. Assume these investors extend fixed-rate, prepayable mortgages to households and also own some noncallable bonds (e.g., Treasuries): the same arbitrageurs are marginal buyers in both MBS and Treasury markets. The key idea is that these arbitrageurs are specialists who are heavily exposed to risks specific to the bond market. The limited risk-bearing capacity of these investors ensures that, at least in the short term, demand curves for aggregate bond risk factors slope downward. In particular, the required return on long-term bonds must rise relative to short-term bonds to induce investors to bear more interest rate risk.

I assume that, at each date t, arbitrageurs have mean-variance preferences over wealth at $t+1$, with possibly time-varying risk aversion, γ_t. This is the discrete-time analog of Vayanos and Vila (2009), who work in continuous time. The key idea is that arbitrageurs are concerned with their interim wealth and, hence, with interest rate risk.They risk having to sell long-term bonds at a capital loss if interest rates rise tomorrow.

Let r_{t+1} denote the log return on one-period riskless bonds between t and $t+1$. By definition $r_{t+1} = y_t^{(1)}$, i.e., the yield on one-period bonds at time t. I assume an exogenous random process for the short rate. One can think of this as being pinned down by monetary policy or by a stochastic short-term storage technology that is in perfectly elastic supply. Thus, from the standpoint of arbitrageurs at time t, the only relevant uncertainty is about $r_{t+2} = y_{t+1}^{(1)}$. The log return on two-period bonds from t to $t+1$ is $r_{t+1}^{(2)} = 2y_t^{(2)} - y_{t+1}^{(1)} = 2y_t^{(2)} - r_{t+2}$, so the excess return on two-period bonds relative to one-period bonds is $rx_{t+1}^{(2)} = r_{t+1}^{(2)} - r_{t+1} = 2y_t^{(2)} - r_{t+1} - r_{t+2}$. Arbitrageurs choose their long-term bond holdings b_t to solve

$$
\begin{aligned}
& \max_{b_t} \left\{ b_t \cdot E_t[rx_{t+1}^{(2)}] - b_t^2 \cdot (\gamma_t/2)\, Var_t[rx_{t+1}^{(2)}] \right\} \qquad (1) \\
= \; & \max_{b_t} \left\{ b_t \cdot (2y_t^{(2)} - r_{t+1} - E_t\left[r_{t+2}\right]) - b_t^2 \cdot (\gamma_t/2)\, Var_t\left[r_{t+2}\right] \right\}.
\end{aligned}
$$

I assume that r_{t+1} and $\gamma_t Var_t[r_{t+2}]$ follow exogenous stochastic processes. Thus,

$$b_t(y_t^{(2)}) = (\gamma_t Var_t[r_{t+2}])^{-1} \cdot (2y_t^{(2)} - r_{t+1} - E_t[r_{t+2}]). \quad (2)$$

I model the MBS positive-feedback channel in a simple way. Let Q_t denote the total dollar quantity of duration risk that needs to be held by fixed income arbitrageurs. In practical terms, I associate Q_t with the total number of ten-year Treasury duration equivalents summed across all US fixed income markets. I allow Q_t to depend on $y_t^{(2)}$. I assume that $Q_t = Q_t(y_t^{(2)})$ with $Q_t'(y_t^{(2)}) > 0$ to capture the fact that MBS duration is increasing in long-term yields. As discussed above, $Q_t'(y_t^{(2)}) > 0$ is a property that holds locally due to time variation in risk sharing between household mortgage borrowers and bond investors. And the vast size of the MBS market relative to the broader US bond market ensures that shifts in MBS duration have a significant effect on aggregate bond market duration.

The equilibrium long-term yield must clear the market at each date, $b_t(y_t^{(2)}) = Q_t(y_t^{(2)})$, which delivers the fixed-point condition

$$y_t^{(2)} = (r_{t+1} + E_t[r_{t+2}])/2 + (\gamma_t/2)\, Var_t[r_{t+2}]\, Q_t(y_t^{(2)}). \quad (3)$$

To ensure that this equilibrium is locally stable, I assume that $1 > (\gamma_t/2)\, Var_t[r_{t+2}]\, Q_t'(y_t^{(2)})$. This ensures that a small exogenous shock to $y_t^{(2)}$ leads to a larger increase in investor demand for bond duration than in duration supply [i.e., $\partial b_t(y_t^{(2)})/\partial y_t^{(2)} > Q_t'(y)$].

The expected excess return on long-term bonds relative to short-term bonds is

$$E[rx_{t+1}^{(2)}] = \gamma_t Var_t[r_{t+2}]\, Q_t(y_t^{(2)}), \quad (4)$$

and the yield spread is

$$y_t^{(2)} - y_t^{(1)} = (E_t[r_{t+2}] - r_{t+1})/2 + (\gamma_t/2)\, Var_t[r_{t+2}]\, Q_t(y_t^{(2)}). \quad (5)$$

Naturally, in the limiting cases in which arbitrageurs are risk neutral ($\gamma_t = 0$) or have to bear no interest rate risk ($Var_t[r_{t+2}]\, Q_t(y_t^{(2)}) = 0$), the expectations hypothesis holds.

2.3. Predictions

The model generates several novel predictions that I now develop in turn.

2.3.1. Forecasting excess bond returns

The model predicts that measures of aggregate bond market duration, which derive most of their power from variation in MBS duration, should positively predict bond excess returns.

Proposition 1 *Both the yield spread and aggregate bond market duration, $Q_t(y_t^{(2)})$, will positively forecast excess bond returns. This remains true in a multivariate forecasting regression so long as $\gamma_t Var_t[r_{t+2}]$ varies over time.*

Why isn't the slope of the yield curve a sufficient statistic for forecasting excess bond returns? In this simple model, the term spread contains information about the expected path of future short rates and term premia, both of which vary over time. Duration measures pertain solely to the latter and, thus, can improve the forecasting power of regressions that include only the term spread. Formally, so long as $\gamma_t Var_t [r_{t+2}]$ varies over time, both the yield spread and duration positively predict excess returns in a multivariate forecasting regression. According to Eq. (4), expected returns equal the product of the time-varying quantity of risk, $Q_t(y_t^{(2)})$, and the time-varying price per unit of risk, $\gamma_t Var_t [r_{t+2}]$. Aggregate duration provides an accurate signal of the quantity of risk. Eq. (5) shows that the yield spread contains information about both the quantity of risk and the price of risk, but it is a noisy indicator of expected returns because it contains an expectations hypothesis component, $(E_t [r_{t+2}] - r_{t+1})/2$. Thus, adding $Q_t(y_t^{(2)})$ to a regression that contains the term spread can raise the forecasting power, particularly if the time series variation in $(E_t [r_{t+2}] - r_{t+1})/2$ is high.

More generally, the aggregate supply of duration contains valuable information about term premia that cannot be easily recovered from the yield curve. For example, duration could help to forecast excess bond returns, even controlling for the term spread, the Cochrane and Piassezi (2005) factor, or other simple yield-based proxies for term premia. Nonetheless, it is not my intention to argue that duration is an unspanned state variable in the sense of Duffee (2011) or Joslin, Priebsch, and Singleton (2013). In classic term structure approaches, if the true model is known, yields can be inverted to obtain the full set of state variables. An unspanned state variable is a variable that is useful for forecasting returns but has (almost) no impact on current yields and cannot be recovered in this way.[4] Specifically, the Internet Appendix shows that I can nest all of my predictions in an affine term structure model similar to Vayanos and Vila (2009) in which MBS duration is a spanned state variable that impacts current yields. Thus, assuming a stationary data generating process, all the information about future bond returns would be contained in current yields, and duration would not add any further information.

In practice, however, the true model generating bond yields is not known. Furthermore, it seems likely that the true model evolves over time due to changes in the macroeconomy, the conduct of monetary policy, the market structure, and the behavior of market participants. For instance, the impact of MBS duration on term premia and bond market spreads has grown over time, arguably because fluctuations in duration have grown larger relative to investor risk tolerance. Thus, in practice, it is natural that duration is a useful summary statistic containing additional information about future excess bond returns.

[4] As Duffee (2011) explains, this would arise if the evolution of short rates under the risk-neutral (pricing) measure is independent of some variable, implying that it has no impact on current yields, even though that variable is relevant for forecasting future short rates under the objective measure. Such a situation could arise if a variable had offsetting effects on the evolution of future short rates and future term premia, e.g., some scary, bad news that raises future expected term premia but lowers future expected short rates.

2.3.2. *Excess sensitivity of long-term rates to short rates*

The negative convexity of the US fixed income market naturally generates excess sensitivity of long-term yields to movements in short-term rates. To see this, consider the static case in which $y^{(2)} = (r_1 + E[r_2])/2 + (\gamma/2)\,Var[r_2]\,Q\left(y^{(2)}\right)$ and consider a change in r_1, holding fixed all other parameters including $E[r_2]$. Because $1 > (\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right) > 0$,

$$\frac{\partial y^{(2)}}{\partial r_1} = \frac{1}{2}\sum_{i=0}^{\infty}\left((\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right)\right)^i = \frac{1/2}{1-(\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right)} > \frac{1}{2}. \tag{6}$$

This excess sensitivity stems from an MBS duration spiral. A small initial shock to the short rate of dr_1 directly raises long-term yields by $(1/2)\,dr_1$ due to the expectations hypothesis. The rise in yields extends the duration of MBS, which raises the term premium by $(1/2)\,[(\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right)]dr_1$; the resulting rise in yields further extends the duration of MBS, which further raises the term premium by $(1/2)\,[(\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right)]^2 dr_1$, and so on.[5]

The same excess sensitivity point can be made in terms of forward rates. Working with forward rates allows for cleaner empirical tests because there is no direct expectations hypothesis term. The forward rate is $f^{(2)} = 2y^{(2)} - r_1 = E[r_2] + \gamma Var[r_2]\,Q(\left(r_1 + f^{(2)}\right)/2)$. Consider a change in r_1, holding fixed all other parameters including $E[r_2]$.[6] Thus,

$$\frac{\partial f^{(2)}}{\partial r_1} = \frac{(\gamma/2)\,Var[r_2]\,Q'(y^{(2)})}{1-(\gamma/2)\,Var[r_2]\,Q'\left(y^{(2)}\right)} > 0. \tag{7}$$

Proposition 2 *If aggregate bond market duration is increasing in yields ($Q'\left(y^{(2)}\right) > 0$), long-term yields and forward rates will be excessively sensitive to movements in short-term rates. Furthermore, both long-term real and nominal rates will exhibit excess sensitivity.*

Thus, the model departs from Vayanos and Vila (2009), who assume that $Q'\left(y^{(2)}\right) < 0$, meaning that arbitrageurs must hold a lower net supply of long-term bonds when long-term yields are high. The assumption that $Q'\left(y^{(2)}\right) < 0$ implies that distant forwards underreact to movements in short rates. However, the bulk of the empirical evidence suggests that distant forwards overreact, not underreact, to movements in short rates.[7]

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[5] Changes in expected future short rates also lower risk premia in the model: $\partial y^{(2)}/\partial E[r_2] = \partial y^{(2)}/\partial r_1 > 1/2$. The MBS convexity effect suggests that Federal Open Market Committee forward guidance could impact term premia. Intuitively, the amplification mechanism is the same for an independent change in either r_1 or $E[r_2]$.

[6] By considering sufficiently distant forwards, it is reasonable to assume that expected short rates in the distant future are insensitive to current short rates, i.e., that $\partial E[r_2]/\partial r_1 \approx 0$. Even if $\partial E[r_2]/\partial r_1 \neq 0$, then $\partial f^{(2)}/\partial r_1 > \partial E[r_2]/\partial r_1$, so the excess sensitivity result still carries through.

[7] Vayanos and Vila (2009) assume that $Q'\left(y^{(2)}\right) < 0$ because it enables their baseline model—featuring an stochastic short rate, no independent supply shocks, and constant arbitrageur risk tolerance—to explain why the term spread forecasts excess bond returns. The intuition is that a rise in short rates, which directly raises long yields and lowers the term spread due to the expectations hypothesis, reduces the amount of long bonds that arbitrageurs must hold and, thus, the term premium. To jointly explain the forecasting power of the term spread and excess sensitivity, one could entertain a more complicated model in which $Q_t'(y_t^{(2)}) > 0$ and, in addition to short rate shocks, there are independent shocks to the supply of duration or arbitrageur risk tolerance.

The model also generates a novel prediction about how the excess sensitivity of long-term forwards varies over time.

Proposition 3 *The excess sensitivity of long-term rates to short rates is more pronounced when the fixed income market is more negatively convex* ($\partial^2 y^{(2)}/\partial r_1 \partial Q'\left(y^{(2)}\right) > 0$).

2.3.3. Excess volatility of long-term interest rates

The model also has implications for the volatility of long-term yields. Suppose it is time t and consider the long-term yield at time $t+1$: $y_{t+1}^{(2)} = (r_{t+2} + E_{t+1}[r_{t+3}])/2 + (\gamma/2) Var[r_{t+3}] Q(y_{t+1}^{(2)})$. Assume the only uncertainty is about $r_{t+2} + E_{t+1}[r_{t+3}]$, so that

$$y_{t+1}^{(2)} - E_t[y_{t+1}^{(2)}] \approx \frac{\partial y^{(2)}}{\partial r_1}|_{y^{(2)}=E_t[y_{t+1}^{(2)}]} \cdot (E_{t+1} - E_t)(r_{t+2} + r_{t+3}). \tag{8}$$

Thus,

$$Var_t[y_{t+1}^{(2)}] \approx \left(\frac{1/2}{1 - (\gamma/2) Var[r_2] Q'(E_t[y_{t+1}^{(2)}])}\right)^2 \cdot Var_t[r_{t+2} + E_{t+1}[r_{t+3}]]. \tag{9}$$

Proposition 4 *All else equal, long-term interest rate volatility,* $Var_t[y_{t+1}^{(2)}]$*, is increasing in the negative convexity of the fixed income market,* $Q'(E_t[y_{t+1}^{(2)}])$.

Excess volatility of long-term rates is a natural corollary of excess sensitivity. However, the negative convexity of the MBS market acts as a positive-feedback loop that amplifies the effects of any primitive shock (e.g., short-term rate, investor risk tolerance, or bond supply) that moves bond yields. Thus, the growth of the MBS market could have led to a secular increase in excess volatility. However, there is an additional time series prediction: Long rates should be particularly volatile at times when the MBS market is most negatively convex. Perli and Sack (2003) and Duarte (2008) develop a similar hypothesis and find support for it in recent US data.

2.3.4. Allowing for multiple bond maturities

In the Internet Appendix, I extend the model with only short- and long-term bonds to allow for multiple bond maturities. The extension is a discrete-time version of the no-arbitrage term structure model developed by Vayanos and Vila (2009) and Greenwood and Vayanos (2014). I add the MBS convexity effect to this model in a simple fashion. I assume that the aggregate supply of duration that must be held by arbitrageurs rises when interest rates rise. However, I assume that these duration shocks dissipate quickly: duration shocks are transitory and are expected to mean-revert over time. These assumptions are a simple way to capture the dynamics of MBS duration explained in the paper.

Solving the model, I obtain a discrete-time affine model of the term structure with two state variables, the current short rate and the current level of MBS duration, which depends on the past path of interest rates. The model generates three predictions.

1. An MBS duration shock raises the current expected excess returns on long-term bonds over short-term bonds. Furthermore, because MBS duration is expected to quickly revert to its long-run mean, this effect is short-lived in expectation.

2. A stronger MBS convexity effect (i.e., a more negatively convex MBS universe) increases the sensitivity of long-term yields and forward rates to movements in short rates.

3. A stronger MBS convexity effect increases the volatility of long-term yields and forwards.

These three results echo those from the simple model. However, the multiple maturity extension delivers two additional predictions that I can take to the data.

4. An MBS duration shock has a larger effect on the expected excess returns on longer-term bonds than those on intermediate-term bonds. This is a natural consequence of the fact that an MBS duration shock raises the current duration risk premium. The expected returns on long duration bonds move more than those on intermediate-duration bonds.

5. However, because they are expected to quickly dissipate, shocks to MBS duration have a humped-shaped effect on the yield curve and the forward rate curve, i.e., a shock to MBS duration increases the curvature of the yield curve. As in Greenwood and Vayanos (2014), the effects of a supply shock on yields, which equals the effect on the bond's average returns over its lifetime, are greater for intermediate-term bonds than for long-term bonds when the supply shock is transient.

These two predictions mean that it is impossible to infer whether there was, say, decline in the duration risk premium simply by asking whether the yield curve flattened. And, notably, the last prediction suggests that transient shocks to MBS duration could account for some of the predictive power of the tent-shaped combination of forward rates identified by Cochrane and Piassezi (2005).

2.3.5. The role of delta-hedging: stock versus flow effects

One important simplification in the current model is that only one class of investors owns MBS and prices interest rate risk. In practice, two sets of intermediaries own MBS. One set of intermediaries delta-hedge the embedded prepayment option and, thus, bear a constant amount of interest rate risk over time. Other investors do not delta-hedge and, thus, bear a time-varying amount of risk.[8]

Does it matter whether some MBS holders delta-hedge the prepayment option? In principle, the answer is "no" because the relevant hedging flows correspond one-for-one with changes in the aggregate quantity of duration risk. To see this, consider a modification of the model featuring a set of banks that extend all mortgages to households. Assume banks do not bear interest rate risk

[8] Conversations with market participants suggest that the government-sponsored enterprises and commercial banks have historically been the most prominent delta-hedgers of MBS.

and instead delta-hedge their interest rate exposure. Specifically, as the prepayment option moves into-the-money, the banks buy noncallable bonds (e.g., Treasuries) to keep the total duration of their assets fixed (and equal to that of their liabilities). Finally, assume that the remaining supply of Treasuries is held by risk-averse bond arbitrageurs, and continue to assume that households do not hedge their time-varying interest rate exposure.

Suppose long-term rates suddenly drop, which raises mortgage refinancing expectations. From the perspective of banks, the existing mortgages now behave like short-term bonds. As a result, banks significantly reduce their delta hedges. That is, the banks buy more long-term Treasuries, financed by selling short-term T-bills, to maintain their previous asset duration. When borrowers refinance, banks unwind these transactions. Thus, banks have a temporarily elevated hedging demand for longer-term bonds. As a result, there is a temporary reduction in the amount of interest rate risk that bond arbitrageurs must bear, so the expected excess return on long-term bonds must fall in equilibrium.[9]

Clearly, the delta-hedging banks do not play an important role in this story. Hedging flows in the model with banks correspond one-for-one to changes in the aggregate quantity of duration risk born by investors in the model without banks. As a result, equilibrium bond prices in the modified model are the same as in the simpler model that omitted banks.[10]

3. Empirical analysis

3.1. Measures of MBS duration and convexity

The aggregate duration of the MBS market is not a simple function of the current mortgage rate. MBS duration also depends critically on the distribution of outstanding mortgage coupons and, in this way, reflects the past path of mortgage rates. Furthermore, times arise when aggregate refinancing and thus MBS duration are more or less sensitive to changes in long yields.

Fortunately, the major brokerage firms publish widely followed estimates of the effective duration of the US MBS market. Duration is the semi-elasticity of price with respect to yield [$DUR = -P'(y)/P$ is the percentage change in price for a small change in yield] and is the most widely used measure of interest rate risk. Bond convexity measures the curvature of the price-yield relation: $CONV = (1/2)\,P''(y)/P$. Thus, the percentage change in price following a

[9]Instead of dynamically delta-hedging the interest rate exposure of MBS, investors can statically hedge their exposures by purchasing interest rate options. Regardless of whether MBS investors pursue a dynamic or a static hedging strategy, other investors must take the other side of these trades. Thus, in the aggregate, fixed income investors must bear a time-varying amount of interest rate risk that impacts equilibrium term premia.

[10]There are reasons to think that the extent of hedging could play a nontrivial role in practice. The MBS market might not be perfectly integrated with the broader bond market, and the extent of hedging could reflect the degree of integration. Suppose there are two types of MBS investors: (1) delta-hedging intermediaries (e.g., the government-sponsored enterprises and banks) who maintain a constant asset duration, and (2) sleepers who inelastically buy a fixed quantity of MBS irrespective of the duration risk they are taking (e.g., foreign official holders). The duration that must be absorbed by arbitrageurs is $\theta_t Q_t^{MBS}(y_t^{(2)})$. Thus, the fraction of MBS held by hedgers, an indicator of the degree of market integration, determines the extent to which shifts in MBS duration impact term premia.

Δy parallel shift in the yield curve is

$$\% \left(\Delta P/P\right) \approx -DUR \cdot \Delta y + CONV \cdot \left(\Delta y\right)^2 . \tag{10}$$

The duration of a negatively convex bond rises with yields. By contrast, the duration of a positively convex bond falls with yields.

For callable bonds such as MBS, $P(y)$ is calculated using a stochastic term structure model and a prepayment model that forecasts MBS cash flows as a function of future rates. Using the riskless yield curve and parameters governing interest rate volatility as inputs, one simulates a large number of scenarios for the future path of rates. One computes the expected cash flows in each scenario using the mortgage prepayment model and then discounts these expected cash flows using the implied zero-coupon curve in that scenario. The model-implied price is computed by taking the probability-weighted average of the discounted cash flows in each scenario.[11]

In the baseline results, I use estimates of MBS duration from Barclays Capital, which are available from Datastream beginning in 1989. Aside from data availability, starting the analysis in the late 1980s is sensible because the MBS market rose to prominence only in the mid-1980s. The Barclays (formerly Lehman Brothers) Fixed Income Indices are the most widely followed set of bond indices in the US. However, as shown in the Internet Appendix, I obtain nearly identical results using the duration of the Bank of America (formerly Merrill Lynch) US Mortgage Master index, which is available starting in 1991. The correlation between the Barclays and Bank of America MBS duration measures is 0.79.

The Barclays US MBS index covers mortgage-backed pass-through securities guaranteed by Government National Mortgage Association (Ginnie Mae), Federal National Mortgage Association (Fannie Mae), and Federal Home Loan Mortgage Corporation (Freddie Mac), collectively known as US Agency MBS. The index is composed of pass-through securities backed by conventional fixed-rate mortgages. The MBS index does not include nonagency or private-label MBS (e.g., MBS backed by Jumbo, Alt-A, or subprime mortgages).[12]

I examine three related measures of duration.

1. DUR_t^{AGG} is the effective duration of the Barclays Aggregate Index and measures the percentage change in the US bond market value following a parallel shift in the Treasury yield curve. Barclays Aggregate Index is a proxy for the broad US fixed income market and includes Treasuries, Agency debentures, US Agency MBS, investment-grade corporates, and some asset-backed securities. Many bond portfolios are benchmarked relative to the Barclays Aggregate Index.

[11] To ensure the model-implied price equals the market price, analysts plug their discount function using an option-adjusted spread (OAS), the constant spread one must add to the riskless curve in all scenarios so the model-implied price equals the market price. Effective duration and convexity are computed at the current OAS.

[12] Conventional mortgages satisfy several size, FICO, and loan-to-value (LTV) requirements. For instance, mortgages in government-sponsored enterprise MBS are subject to the conforming loan size limit set by Congress, typically have FICO scores over 620, and have a maximum LTV of 80%. Barclays MBS Index is formed by grouping MBS pools into generic pools based on agency, program (e.g., 30-year or 15-year), mortgage coupon, and origination year. A generic pool is included in the index if it has a contractual maturity greater than one-year and more than $250 million outstanding.

2. DUR_t^{MBS} is the effective duration of the Barclays MBS Index and measures the percentage change in US Agency MBS market value following a shift in the yield curve.

3. $DUR\_CNTRB_t^{MBS} \equiv \left(MV_t^{MBS}/MV_t^{AGG}\right) \cdot DUR_t^{MBS}$ is the *contribution* of MBS to Barclays Aggregate Index duration:

$$DUR_t^{AGG} = \overbrace{\frac{MV_t^{MBS}}{MV_t^{AGG}} \cdot DUR_t^{MBS}}^{\substack{\text{Duration contribution of MBS:} \\ DUR\_CNTRB_t^{MBS}}} + \overbrace{\left(1 - \frac{MV_t^{MBS}}{MV_t^{AGG}}\right) \cdot DUR_t^{OTH}}^{\substack{\text{Duration contribution of other sectors:} \\ DUR\_CNTRB_t^{OTH}}} . \quad (11)$$

Scaling DUR_t^{MBS} this way captures the fact that shifts in MBS duration have had a growing impact on aggregate bond market duration due to the growth of the MBS market. $DUR\_CNTRB_t^{MBS}$ proxies for the transient component of aggregate bond market duration due to MBS and constitutes my preferred forecasting variable.

3.2. Understanding MBS duration

Before proceeding with the analysis, I explain several facts about MBS duration that play a key role in the story. First, due to the embedded refinancing option, the average duration of MBS is low. Second, MBS duration rises as interest rates rise. This negative convexity property underlies the positive-feedback dynamic emphasized in the paper. Third, shocks to the MBS duration are transient, having a half-life of roughly 5.5 months. Finally, shifts in MBS duration are large relative to investors' risk-bearing capacity. In combination, the last two properties mean that shocks to MBS duration generate significant, short-lived shifts in bond risk premium.

3.2.1. Average MBS duration is low

Panel A of Fig. 1 plots my three duration measures over time. Table 1 presents summary statistics for the main variables in the paper. My monthly sample runs from January 1989 to April 2011. DUR_t^{MBS} averages 3.35 years with a minimum of 0.58 years in May 2003 and a maximum of 4.83 years in May 1994. That is, the typical price-yield sensitivity of MBS is roughly equivalent to that of a 3.5 year, zero-coupon bond.

[Insert Fig.1 and Table 1 about here]

The low average duration of MBS reflects two factors. First, the self-amortizing nature of fixed-rate mortgages lowers their duration relative to nonamortizing bonds. Second, and more important, the prepayment option reduces the duration of MBS. A position in a callable bond is equivalent to a position in a similar noncallable bond plus a short position in an interest rate call option to repurchase the bond at par. Over the past 25 years, the delta of the typical prepayment option has been substantial. The option is typically struck slightly out-of-the-money at origination, so the steady decline in rates since the mid-1980s means most prepayment options have remained near the money.[13]

3.2.2. Negative convexity of MBS

MBS are typically negatively convex. Their duration rises as interest rates rise due to a decline in expected prepayments. Empirically, a 100 basis points (bps) increase in the ten-year Treasury yield has been associated with a 1.37-year increase in DUR_t^{MBS} and a 0.34 increase in DUR_t^{AGG} since 1989.[14]

MBS convexity has played a larger role in US bond markets over time for two reasons. First, the MBS market has grown significantly. The MBS index grew from 25% of the Barclays Aggregate Index in 1989 to 40% in 2008 but had fallen to 30% as of early 2012 due to the post-crisis surge in Treasury borrowing. Second, advances in information technology and heightened competition in mortgage banking have reduced the costs of refinancing. As a result, refinancing now responds more aggressively to changes in primary mortgage rates, so the MBS market has become more negatively convex. For instance, prior to 2000, a 100 bps increase in the ten-year yield was associated with a 1.09-year increase in the MBS duration. Since 2000, a 100 bps increase in the ten-year yield was associated with a 1.57-year increase in the MBS duration.[15]

Another way to illustrate the negative convexity of the MBS market is to examine aggregate refinancing behavior. MBS duration is low when expected future refinancing is high. And current refinancing activity is a strong signal of refinancing over the near-term. Panel C of Fig. 1 plots the Mortgage Bankers' Association Refinancing Index, which reflects the raw number of mortgage applications classified as refinancings, versus a measure of aggregate refinancing incentives. Specifically, I show the Hodrick-Prescott (1997) filtered version of the log-refinancing index, essentially the percentage deviation of refinancings from an estimated trend. A borrower has strong incentives to refinance when the primary rate is below her mortgage coupon, so my measure of aggregate refinancing incentives is the difference between the average coupon of MBS in the Barclays MBS Index and the 30-year primary mortgage rate from Freddie Mac's survey, $c_t - y_{M,t}$. As shown in Panel C, a strong positive relation exists between the two series. The corresponding time series regression has a R^2 of 0.65 and suggests that a 100 bps increase in refinancing incentives boosts aggregate refinancing by 85% relative to trend.

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[13] The price of the callable bond is $P_C(y) = P(y) - \mathcal{C}(P(y))$, where $P(y)$ is the price of the underlying noncallable bond and $\mathcal{C}(P)$ is the price of the interest rate call. Because $0 < \mathcal{C}'(P) < 1$,

$$\overset{\text{Duration of callable}}{-P_C'(y)} = \overset{\text{Noncallable duration equivalents}}{(1-\mathcal{C}'(P(y)))} \times \overset{\text{Duration of noncallable}}{-P'(y)} < \overset{\text{Duration of noncallable}}{-P'(y)} ;$$

$$\overset{\text{Convexity of callable}}{P_C''(y)} = \overset{\text{Noncallable duration equivalents}}{(1-\mathcal{C}'(P(y)))} \times \overset{\text{Convexity of noncallable}}{P''(y)} - \overset{\text{Call option }\Gamma>0}{\mathcal{C}''(P(y))} \times \overset{(\text{Duration of noncallable})^2}{(P'(y))^2} .$$

Because $P''(y) > 0$ and $\mathcal{C}''(P) > 0$, the sign of $P_C''(y)$ is ambiguous. However, because callable bonds become more negatively convex as rates decline, $P_C''(y) < 0$ unless the prepayment option is deeply out-of-the-money.

[14] Estimating $\Delta DUR_t^{MBS} = a + b \cdot \Delta y_t^{(10)} + \varepsilon_t$ using monthly differences, I find $b = 1.37$ ($R^2 = 20.29$).

[15] Beginning in 1997, Barclays estimates the effective negative convexity of the MBS index. If I estimate $\Delta DUR_t^{MBS} = a + b \cdot NCONV_t^{MBS} \times \Delta y_t^{(10)} + \varepsilon_t$, I obtain $a = 0.02$ and $b = 0.97$ ($R^2 = 0.75$) precisely as one would expect. Furthermore, the estimated negative convexity has trended up over time.

3.2.3. Shocks to MBS duration are transient

Shocks to the bond market duration are transient. Specifically, the one-month autocorrelation of the duration of the Barclays Aggregate Index, DUR_t^{AGG}, is 0.88. The transient nature of aggregate duration dynamics is driven by the MBS component of the index. Specifically, the one-month autocorrelation of DUR_t^{MBS} is also 0.88, which implies that MBS duration shocks have a half-life of 5.5 [$=\ln(0.5)/\ln(0.88)$] months. By contrast, the duration of the non-MBS portion of the index is far more persistent with a one-month autocorrelation of 0.96, implying a half-life for non-MBS duration shocks of 17.5 months.

Why are shocks to MBS duration so transient? In the case of a downward shock to rates, the explanation is fairly mechanical. A decline in rates raises expected prepayments, thus lowering the amount of duration risk that holders of outstanding MBS face over the near-term. As homeowners respond to this interest rate shock, prepayments rise. When a borrower prepays, she replaces a high coupon bond with a deep in-the-money prepayment option with a lower coupon bond with a slightly out-of-the-money prepayment option. Borrowers effectively restrike the interest rate call options embedded in their mortgages. The net effect of this refinancing transaction raises the total quantity of duration that investors need to bear. However, because prepayments are gradual, there are temporary declines in aggregate MBS duration, which then predictably revert over time. By contrast, in the case of an upward shock to rates, there is no expectation of such rapid mean reversion in MBS duration. This is because upward shocks to rates are expected to slow refinancing activity, so there is no mechanical restriking effect.

3.2.4. Shifts in MBS duration are large relative to risk-bearing capacity

Due to the vast size of the MBS market, shifts in MBS duration have a large effect on the aggregate amount of interest rate risk born by investors. And shifts in MBS duration drive almost all high-frequency variation in aggregate bond market duration. This is shown in Panel B of Fig. 1 which uses Eq. (11) to decompose 12-month changes in aggregate bond market duration into an MBS component and a non-MBS component. The MBS component accounts for the vast bulk of the variation in aggregate duration. The R^2 from a regression of DUR_t^{AGG} on $DUR\_CNTRB_t^{MBS}$ is 0.61 in levels, 0.81 in 12-month changes, and 0.91 in one-month changes.

To get a sense of the dollar magnitudes, investors convert aggregate duration statistics into ten-year US Treasury equivalents. If the market value of some asset class X is Q_X with effective duration D_X and the duration of a ten-year US Treasury is D_{T10}, then X represents $Q_X \cdot (D_X/D_{T10})$ ten-year Treasury equivalents. Panel D of Fig. 1 plots the detrended component of MBS duration in ten-year Treasury equivalents [the trend is estimated using the Hodrick-Prescott (1997) filter]. I convert historical dollars to 2012 dollars using the consumer price index. The figure shows that the resulting shifts in duration supply are massive. In 26 sample months, detrended MBS ten-year equivalents exceed +\$500 billion or −\$500 billion.

These are arguably very large shifts in the quantity of duration risk relative to investors' risk-bearing capacity. I can compare the shifts in MBS duration with the Fed's Quantitative Easing operations. From 2008Q4 to 2010Q1, the Fed purchased $\$1,250$ billion MBS with an effective duration of roughly three years. Because the duration of the ten-year Treasury note was roughly eight years at this time, this amounted to a reduction of \$469 billion [$\$469 = \$1,250 \times (3/8)$] ten-year equivalents of duration that needed to be held by investors (assuming no further change in interest rates). Thus, Panel D shows that numerous shifts in MBS duration since the late 1980s have been of a comparable scale to the Fed's QE operations. Furthermore, these shifts can occur rapidly, which is likely to be relevant if capital is slow-moving.

3.3. Forecasting excess bond returns

I now show that measures of MBS duration are strong predictors of excess government bond returns.

3.3.1. Basic forecasting results

I use data from Gürkaynak, Sack, and Wright (2007) on the nominal Treasury yield curve as updated regularly by the Federal Reserve Board. The log excess return on an n-year zero-coupon bonds is defined as $rx_{t+1}^{(n)} = n \cdot y_t^{(n)} - (n-1) \cdot y_{t+1}^{(n-1)} - y_t^{(1)}$. Panel A of Table 1 provides summary statistics on 12-month excess returns ($rx_{t+1}^{(n)}$) and yield spreads ($y_t^{(n)} - y_t^{(1)}$) for $n = 5, 10,$ and 20. I also summarize information on the instantaneous forward rates $f_t^{(m)}$ for $m = 1, 2, 3, 4, 5$.

Table 2 uses the three bond duration measures to forecast excess returns on ten-year nominal US Treasuries. Specifically, Table 2 presents forecasting regressions of the form

$$rx_{t+1}^{(10)} = a + b \cdot DUR_t + \mathbf{c}'\mathbf{x}_t + \varepsilon_{t+1}^{(10)}. \tag{12}$$

For the sake of comparability with the recent literature, the regressions are estimated with monthly data, so each month I am forecasting excess returns over the following 12 months. To deal with the overlapping nature of returns, t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. I estimate these regressions with and without other forecasting variables identified in the literature on bond risk premia. Specifically, I control for the term spread following Campbell and Shiller (1991) and the first five forward rates following Cochrane and Piazzesi (2005). I obtain similar results controlling for the ten-year forward rate spread following Fama and Bliss (1987) or if I control for longer-dated (e.g., seven-, ten-, and 20-year) forwards as in Cieslak and Povala (2013).

[Insert Table 2 about here]

Table 2 shows that aggregate duration (DUR_t^{AGG}), MBS duration (DUR_t^{MBS}), and MBS duration contribution ($DUR\_CNTRB_t^{MBS}$) are each very strong predictors of bond excess returns. MBS duration contribution is the strongest predictor, both in univariate specifications and in the multivariate regressions. Duration adds incremental forecasting power over and above the term spread, the forward rate spread, and arbitrary combinations of forward rates. Specifically, as shown in Table 2, duration significantly raises the R^2 in these specifications, and the R^2s are already high. For instance, Column 2 shows that using the first five forward rates as predictors delivers a forecasting R^2 of 0.31. If MBS duration contribution is added to this multivariate specification as in Column 11, the forecasting R^2 jumps to 0.46.

Comparing Columns 2 and 11, the inclusion of duration measures roughly halves the absolute magnitude of the coefficients on forward rates in the augmented Cochrane and Piazzesi (2005) regression. I return to this observation below when I explain why MBS duration could help to account for some of the forecasting power of Cochrane and Piazzesi's tent factor.

[Insert Fig. 2 about here]

Fig. 2 illustrates the basic forecasting result. Specifically, Panel A shows a scatter plot of 12-month future excess returns versus MBS duration contribution. These plots correspond to the multivariate estimates in Column 10. Specifically, I plot the component of excess returns that is orthogonal to the term spread versus the component of duration that is orthogonal to the term spread. The strong positive relation is evident in Panel A. Panel B plots 12-month future excess returns versus initial MBS duration contribution over time. The figure shows that MBS duration adds significant forecasting power because duration extensions in 1991, 1994, 1999, 2004, 2007, 2008, and 2011 were each followed by high bond excess returns. Conversely, duration contractions in 1993, 1995, 1998, 2002, 2003, and 2009 were followed by low excess returns. Thus, Fig. 2 shows that the basic forecasting result is not driven by one subsample or by a handful of outlying events. In untabulated regressions, I obtain strong results when I separately examine the 1989–1999 subsample or the 2000–2011 subsample.

What is the economic magnitude of the estimated effects? The coefficient in Column 10 indicates that a one-year increase in $DUR\_CNTRB_t^{MBS}$ raises expected excess returns on ten-year bonds by 14.629% over the following 12 months. A 1 standard deviation move in MBS duration contribution is 0.29 years, so this implies that a 1 standard deviation increase in duration raises expected excess returns by $4.20\% = 14.629 \times 0.29$. Assuming this shift has no effect on expected returns beyond 12 months (analysis below suggests this is a reasonable assumption), this corresponds to a rise in ten-year yields of 42 [= 420/10] bps today. Thus, the estimated effects are highly economically significant. However, given the size and speed of the MBS duration shocks relative to the scale of arbitrage capital, the effects do not seem implausibly large. I return to this point below when I compare the implied price impact from these regressions with estimates from the literature evaluating the Federal Reserve's Quantitative Easing policies.

These forecasting regressions suggest a simple market-timing strategy for diversified investors (e.g., endowments and pensions) who allocate capital between the bond market and other financial markets. Specifically, diversified investors should overweight long-term bonds when MBS duration is high and underweight long-term bonds when MBS duration is low. Unlike the specialist bond arbitrageurs emphasized in the model, the changes in the aggregate quantity of interest rate risk perceived by these investors is miniscule because bonds are only a small portion of their overall financial portfolios. Thus, such diversified investors perceive the resulting time-variation in expected returns as time–varying alpha as opposed to a time-varying bond risk premium.

Naturally, the t-statistics from the forecasting regressions can be translated directly into a statement about the Sharpe ratio of strategy that buys (sells) long-term bonds to the extent that $DUR\_CNTRB_t^{MBS}$ is above (below) its unconditional average.[16] Estimating a univariate regression with quarterly data yields a t-statistic of 2.91, implying a strategy Sharpe ratio of $0.303 = 2.91 \div \sqrt{92}$ per quarter. Assuming that strategy returns are independent over time, this implies an annual Sharpe ratio of $0.606 = \sqrt{4} \times 0.303$. The fact that this timing strategy appears to be so profitable suggests that diversified investors do not fully exploit it. This could be because various frictions combine to limit the speed at which arbitrage capital flows across markets as in Duffie (2010). Or this could simply be because many diversified investors are unaware of the high Sharpe ratio offered by the strategy.

3.3.2. MBS duration accounts for all the forecasting power of aggregate duration

Table 3 shows that all of the forecasting power of DUR^{AGG} derives from the MBS component. Specifically, I decompose DUR^{AGG} into the sum of $DUR\_CNTRB_t^{MBS}$ and $DUR\_CNTRB_t^{OTH}$ as in Eq. (11). The transient MBS component accounts for the vast majority of all high-frequency variation in aggregate bond market duration. The 12-month autocorrelations of DUR^{AGG}, $DUR\_CNTRB_t^{MBS}$, and $DUR\_CNTRB_t^{OTH}$ are 0.16, 0.04, and 0.64, respectively. Thus, perhaps not surprisingly, the horse races in Columns 4, 8, and 12 of Table 3 show that all of the near-term forecasting power in aggregate bond market duration is attributable to MBS duration. A similar conclusion holds if, instead of $DUR\_CNTRB_t^{OTH}$, I focus more narrowly on the duration of the Barclays Treasury Index.

[Insert Table 3 about here]

In summary, both the duration of the Barclays Aggregate Index and the Barclays MBS Index positively forecasts excess bond returns over the following 12 months. And the duration contribution of MBS, which scales MBS duration relative to the broader market, is the strongest forecaster. However, the measured duration of the Treasury and other non–MBS fixed income markets do not reliably forecast bond returns at the quarterly or annual frequencies considered here, at least not in my 1989–present sample. This is perhaps not surprising if one considers a preferred-habit model with slow-moving capital. In such a model, capital flows in response to persistent changes in bond risk premia, which neutralizes some, but not all, of the effects of persistent shocks to duration supply on expected returns. Thus, from a medium-frequency return forecasting perspective, it could be particularly useful to isolate the transient component of shocks to bond duration supply. And the preceding analysis suggests that MBS duration is an excellent proxy for this transient component of duration supply.[17]

[16] Suppose $r_{t+1} = \alpha + \beta \cdot x_t + \varepsilon_{t+1}$ and consider returns on the strategy $w_{t+1} = r_{t+1}(x_t - E[x_t])$. Then, $SR[w_{t+1}] = \beta\sigma^2[x_t] \div \sqrt{\sigma^2[\varepsilon_{t+1}]\sigma^2[x_t]} = t[b_{OLS}] \div \sqrt{T}$, where $t[b_{OLS}] = \beta \div \sqrt{T^{-1}\sigma^2[\varepsilon_{t+1}](\sigma^2[x_t])^{-1}}$ is the population t-statistic for the ordinary least squares estimator of β.

[17] Admittedly, because my measures of MBS duration reflect forward-looking expectations of future MBS duration but my measures of Treasury duration are not forward-looking, this horse race between MBS and non-MBS

3.3.3. Time signature of MBS duration effects

Variation in MBS duration is expected to be associated with relatively high frequency variation in expected returns for several reasons. First, if current MBS duration positively predicts bond excess returns over the following instant, the current duration would not be expected to reliably predict returns on 12-month forward basis because MBS duration is not itself very persistent. Second, slow-moving capital effects as in Duffie (2010) could exist, which implies that the short-run demand curve for duration risk is more inelastic than the long-run demand curve. Specifically, suppose that, in the short run, interest rate risk is priced by specialized bond market investors with a fixed risk tolerance. However, over longer horizons, diversified investors could allocate financial capital between the bond market and other financial markets as in Duffie and Strulovici (2012). Thus, following a positive shock to duration supply, bond risk premia could be expected to jump. The anticipation of large future bond returns would draw in capital from other markets, raising the risk tolerance of bond investors and, thus, reducing the impact of supply on bond risk premia over time in the case of a permanent supply shock.

To formally explore the time signature of my main forecasting result, I use MBS duration to forecast quarterly returns from one to eight quarters ahead. Fig. 3 presents quarterly forecasting regressions of the form

$$rx^{(10)}_{t+(j-1)/4\to t+j/4} = a_{(j)} + b_{(j)} \cdot DUR\_CNTRB^{MBS}_t + \mathbf{c}'_{(j)}\mathbf{x}_t + \varepsilon^{(10)}_{t+(j-1)/4\to t++j/4} \tag{13}$$

for $j = 1, ..., 8$. Fig. 3 then plots the coefficients $b_{(j)}$ versus the horizon j. This series of regressions provides a simple nonparametric way to trace out the impulse response of quarterly excess bond returns following a movement in MBS duration. Specifically, Panel A of Fig. 3 plots the coefficients $b_{(j)}$ from univariate regressions, and Panel B plots the coefficients from multivariate specifications that control for the term spread. Fig. 3 shows that the forecasting power of bond duration is largely located in the following two quarters. The effect decays meaningfully from three to four quarters out, and generally little predictive power is evident beyond five quarters.

[Insert Fig. 3 about here]

To investigate the timing of these effects more parametrically, I estimate first-order vector autoregression (VAR) including excess returns, MBS duration, and the term spread using quarterly data:

$$\begin{aligned} rx^{(10)}_{t+1/4} &= a_1 + b_1 \cdot rx_t + c_1 \cdot DUR_t + d_1 \cdot (y^{(10)}_t - y^{(1)}_t) + \varepsilon_{1,t+1/4}, \\ DUR_{t+1/4} &= a_2 + b_2 \cdot rx_t + c_2 \cdot DUR_t + d_2 \cdot (y^{(10)}_t - y^{(1)}_t) + \varepsilon_{2,t+1/4}, \\ y^{(10)}_{t+1/3} - y^{(1)}_{t+1/4} &= a_3 + b_3 \cdot rx_t + c_3 \cdot DUR_t + d_3 \cdot (y^{(10)}_t - y^{(1)}_t) + \varepsilon_{3,t+1/4}. \end{aligned} \tag{14}$$

Using $DUR\_CNTRB^{MBS}_t = \left(MV^{MBS}_t / MV^{AGG}_t\right) \cdot DUR^{MBS}_t$ as the measure of MBS duration, Fig. 4 shows the resulting simple impulse response of quarterly excess bond returns to a shock

duration is not a completely fair fight.

to MBS duration. As expected, the effect on excess bond returns from a shock to MBS duration is concentrated in the next four quarters. Furthermore, using a parametric VAR, paints a very similar picture of the timing of the effects to the nonparametric approach show in Fig. 3.

[Insert Fig. 4 about here]

3.3.4. Results for multiple maturities

If changes in DUR_t shift the duration risk premium in US bond markets, this would be expected to have a larger effect on the expected returns of long-term bonds than on intermediate bonds. For instance, the coefficient on DUR_t when forecasting 20-year excess bond returns should be larger than when forecasting ten-year excess bond returns. This is illustrated in the left-most plots in Fig. 5 where I estimate

$$rx_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)} \tag{15}$$

for $n = 2, ...20$. I then plot the coefficients, $b_{(n)}$, versus maturity n. I show results for DUR_t^{MBS} in Panel A and $DUR\_CNTRB_t^{MBS}$ in Panel B. As shown in Fig. 5, the coefficients $b_{(n)}$ are increasing in maturity n, consistent with the idea that shifts in MBS duration impact the duration risk premia in bond markets.

[Insert Fig. 5 about here]

How do shifts in MBS duration impact bond yields and the current shape of the yield curve? Due to their transitory nature, MBS duration supply shocks would be expected to have a hump-shaped effect on the yield curve and forward rate curve. Although a transitory rise in the duration risk premium has the largest impact on the price and current expected excess returns of long-term bonds, it has the largest impact on the yield of intermediate-term bonds. As in Greenwood and Vayanos (2014), the intuition is that the impact bond yields equals the effect on a bond's average returns over its lifetime. As a result, a temporary rise in duration risk premium has a greater impact intermediate-term yields than on long-term yields when the supply shock is expected to be short-lived.

These results are shown in the middle and right-most plots in Fig. 5. Specifically, I plot the slope coefficients versus maturity from estimating

$$y_t^{(n)} - y_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)} \tag{16}$$

and

$$f_t^{(n)} - f_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)} \tag{17}$$

for $n = 0, 1, 2, ..., 20$. In other words, I use DUR_t to forecast the decline in yields and forwards over the following 12 months for each maturity. Fig. 5 shows that high level of current MBS duration has a humped shaped effect on the yield and forward rate curves. Specifically, the expected decline in rates over the following 12 months is largest for intermediate yields and forwards, having its maximal effect on two-year yields.

This suggests that transient shocks to MBS duration could account for some of the predictive power of the tent-shaped combination of forward rates identified by Cochrane and Piassezi (2005). The Cochrane and Piazzesi (2005) factor picks up time variation in the curvature of the yield curve and is useful for forecasting transitory variation in bond returns, i.e., variation at frequencies higher than a standard business cycle frequency. Thus, the variation in MBS duration could help explain why a tent-shaped combination of forwards can explain transitory variation in bond risk premia. Of course, any fast-moving state variable that affects bond risk premia should have a hump-shaped effect on yields and forwards as basic matter of no-arbitrage bond pricing logic. However, the literature has generally struggled to produce economically plausible state variables that might induce high-frequency variation in bond risk premium. I argue that MBS duration is one such variable and, thus, the fact that it fights with the Cochrane and Piazzesi (2005) factor is natural.

In summary, the findings for multiple maturities are highly consistent with the model I develop in the Internet Appendix and with prior work by Vayanos and Vila (2009) and Greenwood and Vayanos (2014), who argue that shifts in the maturity structure of borrowing (holding constant total dollar borrowing) can significantly impact term premia. And this indicates that the Fed's Quantitative Easing policies can impact the duration risk premium through a broad portfolio balance channel, consistent with the findings of Gagnon, Raskin, Remache, and Sack (2011) and Greenwood and Vayanos (2014).

3.3.5. Robustness checks

Running a regression of returns on quantities does not, in general, allow one to cleanly identify a demand curve. Instead, to nail down the MBS duration supply channel, I would like to show that some component of MBS duration that is exogenous with respect to interest rates positively forecasts excess bond returns. Otherwise, one might worry that MBS duration is simply correlated with some omitted variable that is driving demand for bonds. Unfortunately, constructing a valid and powerful instrument for MBS duration is difficult. However, because the issues addressed here are of first-rder importance for understanding of bond market dynamics and the determination of risk premia more generally, evidence that is admittedly somewhat indirect should be acceptable. Specifically, I am forced to rely on predictive regressions that provide indirect evidence consistent with the MBS hedging story.

In the Internet Appendix, I address several concerns raised by this indirect approach. First, one could be concerned that variation in DUR_t does itself not drive the term premium but is simply correlated with an omitted variable that does. The Internet Appendix shows that the coefficient on DUR_t is highly robust to the inclusion of additional controls that are thought to be associated with term premia. This should help to alleviate most natural concerns about omitted variable bias. Second, I address the standard econometric concerns that arise when estimating time series forecasting regressions. These econometric issues are not a significant concern in this context. Finally, one could worry that the results are somehow specific to the Barclays indices I use throughout. The Internet Appendix shows that similar forecasting results obtain when using the Bank of America MBS indices instead of the Barclays indices.

3.3.6. Comparison with prior estimates of the price impact of duration supply shocks

I now compare the price-impact effects implied by my regressions with estimates from the recent literature evaluating the Federal Reserve's Large-Scale Asset Purchase (LSAP) programs. While there are competing interpretations (Krishnamurthy and Vissing-Jorgensen, 2013), to draw this comparison, I assume that LSAP's primarily work by impacting term premia. To be clear, I am not arguing that mortgage refinancing explains why LSAP policies do or do not work. However, both mortgage refinancing waves and LSAPs are sources of duration supply shocks that have the potential to impact term premia. Thus, it is interesting to compare the price impact of supply shocks from LSAPs and mortgage refinancing waves.

Different price-impact magnitudes for LSAP- and mortgage refinancing-induced duration shocks are expected for several reasons. First, investor risk-bearing capacity likely varies over time, so larger effects for LSAPs should be expected to the extent they occurred when risk-bearing capacity was low. Further, the LSAP-induced shocks would have been expected to last longer than mortgage duration shocks, suggesting that LSAPs should have a larger impact on yields due to their more persistent effect on term premia. However, LSAPs were arguably far more advertised than mortgage duration shocks, suggesting that mortgage shocks could have a larger impact than LSAPs if, as in Duffie (2010), investors were more inattentive in the former case. Thus, it is not obvious whether larger effects should be expected for MBS duration shocks or for LSAPs.

Table 4 compares the estimated effect of supply shocks measured in ten-year Treasury duration equivalents. Table 4 provides more detail on the LSAPs, estimates of their impact on ten-year yields from the prior literature, and the estimated effects implied by the MBS duration effect. I use my estimates to predict the effect of LSAP announcements on ten-year Treasury yields. Specifically, using specifications of the form $y_t^{(10)} - y_{t+1}^{(10)} = a + b \cdot DUR_t + \varepsilon_{t+1}^{(10)}$ as in Fig. 5, I estimate that a 1 standard deviation decline in DUR_t^{MBS} has lowered ten-year Treasury yields by 36 bps over my 1989-present sample. Based on the relation between DUR_t^{MBS} and MBS ten-year equivalents in Fig. 1, I estimate that a 1 standard deviation decline in DUR_t^{MBS} corresponds to a $503 billion reduction in ten-year equivalents in recent years.[18]

[Insert Table 4 about here]

The first LSAP was announced in late 2008 and early 2009 amidst unusually strained market conditions (Krishnamurthy and Vissing-Jorgensen, 2013). Because LSAP1 represented a cumulative reduction in $750 billion ten-year equivalents, my estimates imply a 54 bps [= 36 bps × (750/503)] decline in ten-year Treasury yields. This implied effect is smaller than event

[18] I obtain this estimate by regressing the deviation of Q_t^{MBS} $(DUR_t^{MBS}/DUR_t^{T10})$ from trend (in constant 2012 dollar) on DUR_t^{MBS} from 2008 to the present.

study-based estimates of the impact of LSAP1 on ten-year yields, which average nearly 100 bps. However, this difference makes sense because intermediary risk-bearing capacity was limited at the time of LSAP1. Table 4 shows that the implied effect is slightly larger for univariate regressions using $DUR\_CNTRB_t^{MBS}$ or for multivariate regressions that control for the term spread. For LSAP2 and LSAP3, which were announced in late 2010 and late 2011, respectively, the effects implied by my estimates are in line with the estimated LSAP announcement effects in the prior literature.

In summary, the price impact of supply implied by my regressions is broadly in line with those found in the recent event-study literature on LSAPs. To be clear, my results do not show that LSAPs have lowered the duration risk premium in bond markets or that they work at all.[19] However, my evidence does provide an out-of-sample proof of concept for LSAP policies that seek to impact market-wide term premia.

3.3.7. A decomposition of aggregate MBS duration

I now decompose aggregate MBS duration. The decomposition shows that DUR_t^{MBS} reflects aggregate refinancing incentives, prepayment burnout, and several other factors. I then show that both aggregate refinancing incentives and these other terms appear to contribute to the forecasting power of DUR_t^{MBS}.

Let $y_{M,t}$ denote the primary mortgage rate, and take first-order expansion of the duration of MBS class i (e.g., MBS backed by Fannie Mae guaranteed 30-year mortgages with a 6% coupon that were originated three years ago) about its coupon c_{it}:

$$DUR_{it} = D_{it} - CONV_{it} \cdot (y_{M,t} - c_{it}) + \epsilon_{it}, \tag{18}$$

where ϵ_{it} is the approximation error. Letting $X_t = \sum_i w_{it} X_{it}$ and $\sigma_t [X_{it}, Y_{it}] = \sum_i w_{it} (X_{it} - X_t)(Y_{it} - Y_t)$ denote the market-weighted average and covariance, respectively, aggregate MBS duration is given by

$$DUR_t = D_t + CONV_t \times (c_t - y_{M,t}) + \sigma_t [CONV_{it}, c_{it}] + \epsilon_t. \tag{19}$$

There are two main terms[20]:

1. *Aggregate refinancing incentives.* The first term depends on the aggregate refinancing incentive, $(c_t - y_{M,t})$. This term can be further decomposed using deviations from time series

[19] Besides moving the duration risk premium, LSAP policies could impact yields throughout the bond market through a variety of distinct channels. This is a lively area of scholarly debate. See Krishnamurthy and Vissing-Jorgensen (2011, 2013). Furthermore, LSAPs could have local effects. Comparing securities with similar durations, D'Amico, English, Lopez-Salido, and Nelson (2012) and Cahill, D'Amico, Li, and Sears (2013) find larger intraday declines in yields on those securities that the Federal Reserve is purchasing.

[20] Further insight into the residual ϵ_t can be gained using a second-order Taylor expansion of DUR_{it}. This suggests that aggregate MBS duration is decreasing in the cross-sectional variation in mortgage outstanding coupons.

average of D_t and $CONV_t$:

$$D_t + CONV_t \times (c_t - y_{M,t}) = [D + CONV \times (c_t - y_{M,t})] + [(D_t - D) + (CONV_t - CONV) \times (c_t - y_{M,t})]. \quad (20)$$

Because D_t and $CONV_t$ likely vary only slowly over time, this suggests that $(c_t - y_{M,t})$ should be a strong linear predictor of excess returns.

2. *Aggregate prepayment burnout*. $\sigma_t [CONV_{it}, c_{it}]$ measures the extent to which low callability is concentrated in high coupon MBS. Aggregate MBS duration is high when aggregate burnout is high. For instance, some analysts argue that burnout has been elevated in recent years because borrowers who are still in higher coupon mortgages cannot refinance due to negative home equity or tighter mortgage underwriting standards.

This works largely as one might expect in the data. Specifically, a strong negative relation exists between DUR_t^{MBS} and $(c_t - y_{M,t})$:

$$DUR_t^{MBS} = \underset{[18.22]}{3.095} - \underset{[-5.56]}{0.952} \cdot (c_t - y_{M,t}), \; R^2 = 0.48. \quad (21)$$

Adding a time trend (t) and interactions terms, I obtain

$$DUR_t^{MBS} = \underset{[16.76]}{3.610} - \underset{[-5.18]}{0.835} \cdot (c_t - y_{M,t}) - \underset{[-0.58]}{0.276} \cdot (c_t - y_{M,t}) \times t - \underset{[-2.01]}{1.041} \cdot t, \; R^2 = 0.58. \quad (22)$$

Thus, MBS duration has trended down by a little more than one year since 1989, and weak evidence exists that $CONV_t$ has become slightly more negative over time.

Consistent with the previous decomposition, Columns 1–3 of Table 5 show that the aggregate refinancing incentive, $c_t - y_{M,t}$, is a strong negative predictor of excess bond returns. My measure of aggregate refinancing incentives is the weighted-average coupon on MBS in the Barclays index minus the current 30-year primary mortgage rate from Freddie Mac. The forecasting power of this variable derives from comparing the primary mortgage rate (a variable that is almost a perfect linear combination of forwards) with the average coupon on outstanding MBS (a variable not spanned by forwards). Naturally, c_t is itself a very slow moving variable that reflects historical path dependencies. Columns 7–9 of Table 5 show that I also obtain similar results if I forecast returns using a normalized version of the Mortgage Banking Association's refinancing index, $REFI_t$. Specifically, the Hodrick-Prescott (1997) filtered version of the log-refinancing index is a strong negative predictor of returns. Thus, Table 5 shows that my main findings are not an artifact of using MBS duration, an admittedly complex model-based construct, and supports the broader hypothesis that aggregate mortgage refinancing activity plays a key role in driving bond risk premia.

[Insert Table 5 about here]

Finally, Columns 4–6 and 10–12 of Table 5 show forecasting horse races between my preferred measure of MBS duration ($DUR\_CNTRB_t^{MBS}$) and refinancing incentives ($c_t - y_{M,t}$) and refinancing activity ($REFI_t$), respectively. Consistent with the previous decomposition, these horse races suggest that much of the information that $DUR\_CNTRB_t^{MBS}$ contains about future excess returns is also contained in $c_t - y_{M,t}$ (and $REFI_t$). However, the horse races suggest that model-based MBS duration estimates do contain additional information above and beyond these simpler measures. This is not surprising as MBS duration reflects the forward-looking supply expectations of market participants. Furthermore, $DUR\_CNTRB_t^{MBS}$ captures effects, such as aggregate prepayment burnout, that are not reflected in the simpler measures.

3.3.8. The growing impact of MBS duration on term premia

Have shifts in MBS duration become a more or less important driver of term premia over time? My story suggests that these effects should have grown far more important. The rapid growth of the MBS market means that these duration shocks have become far more significant relative to the risk tolerance of bond investors. The significant growth of the MBS market relative to the rest of the bond market is illustrated in Panel A of Fig. 6.

[Insert Fig. 6 about here]

Fortunately, I can investigate the evolution of these effects by examining data prior to 1989. Although I have data on the model-based effective duration of MBS beginning only in 1989, Barclays publishes a cruder duration measure called Macaulay duration-to-worst dating back to 1976.[21] I use dur to denote the Macaulay analogs of the more commonly used modified duration measures (DUR). Thus, going back to 1976, I can ask whether dur_t^{MBS} and $dur\_cntrb_t^{MBS} = \left(MV_t^{MBS}/MV_t^{AGG}\right) dur_t^{MBS}$ have predictive power for excess bond returns. While the model-based effective duration measures provide a more accurate measure of the sensitivity of MBS prices to changes in yields, Macaulay duration-to-worst contains similar information in the time series. These two measures are shown in Panel B of Fig. 6. For instance, from 1989 to the present when both measures are available, the correlation between effective duration and Macaulay duration-to-worst of MBS is 0.91 in levels and 0.91 in 12-month changes. I can also construct my measure of aggregate refinancing incentives, i.e., the difference between the average MBS coupon and current primary mortgage rates ($c_t - y_{M,t}$) going back to 1976.

Because Panel A of Fig. 6 shows that the MBS market grew significantly relative to the rest of the bond market from 1976 to the early 1990s, examining the forecasting power of MBS duration over this early period is useful. Table 6 shows subsample forecasting results. I find that the three measures of MBS duration have no predictive power for excess bond returns over the 1976-1988 subsample, but they have significant predictive power from 1989 to the present

[21] Macaulay duration-to-worst is the Macaulay duration corresponding to the yield-to-worst. Specifically, one assumes a deterministic set of future MBS cash flows based on an assumed prepayment profile (typically the expected prepayment profile, which depends on current interest rates). The yield-to-worst and Macaulay duration-to-worst are then simply the yield and Macauley duration assuming this set of future bond cash flows.

(including both the 1989-1999 and 2000+ subsamples). Fig. 7 makes a similar point and shows the results from 60-month (20-quarter) rolling regressions of three-month excess returns on MBS duration measures, controlling for the term spread. The results show an unmistakable upward trend in the predictive power of MBS duration. Over time, the coefficients have grown and are being estimated more precisely, so the t-statistics have grown. The fact that these effects have grown so much stronger over time, precisely as I would predict, provides further indirect evidence for my MBS hedging story.

[Insert Table 6 and Fig. 7 about here]

3.3.9. Predicting foreign bond returns

Do shifts in the duration of US mortgage-backed securities impact term premia in foreign bond markets? This question is addressed in Table 7, which investigates whether the duration of US MBS forecasts the excess returns on long-term foreign bonds. What one should expect to find here depends on the degree of integration between the bond markets in different countries. At one extreme, if foreign bond markets were completely segmented from the US bond market, US MBS duration would be expected to have no forecasting power for foreign excess bond returns. At the other extreme, if there were a single integrated global bond market, term premia would be expected to move in lock-step across national bond markets. In between these two extremes, one would expect US MBS duration to have some predictive power for excess foreign bond returns, albeit less than for US excess bond returns.

[Insert Table 7 about here]

Following Campbell (1999, 2003), I work with International Financial Statistics data and compute excess returns on coupon-bearing long-term foreign government bonds over short-term foreign bond using the Shiller, Campbell, and Schoenholtz (1983) approximation.[22] I compute foreign government bond returns for ten developed countries: Belgium, Canada, Denmark, France, Germany, Italy, Japan, Sweden, Switzerland, and the UK. I then estimate univariate forecasting regressions of the form

$$rx_{c,t+1}^{FOR(n)} = a + b \cdot DUR_t + \varepsilon_{c,t+1}^{FOR(n)} \tag{23}$$

and multivariate regressions controlling for the US term spread and foreign country term spread

$$rx_{c,t+1}^{FOR(n)} = a + b \cdot DUR_t + c \cdot (y_{c,t}^{US(n)} - y_t^{US(1)}) + d \cdot (y_{c,t}^{FOR(n)} - y_t^{FOR(1)}) + \varepsilon_{c,t+1}^{FOR(n)}. \tag{24}$$

To assess the differential forecasting power for foreign bond excess returns as compared with US domestic bond returns, for each country I re-run the regressions with $rx_{c,t+1}^{FOR(n)} - rx_{c,t+1}^{US(n)}$ on

[22] This approximation says that log excess return on an n-year coupon bond over the short rate ($y_t^{(1)}$) is $rx_{c,t+1}^{(n)} \approx D_t^{(n)} y_{c,t}^{(n)} - (D_t^{(n)} - 1) y_{c,t+1}^{(n-1)} - y_t^{(1)}$, where $Y_{c,t}^{(n)}$ is the yield on an n-year coupon bond, $y_{c,t}^{(n)} = \ln(1 + Y_{c,t}^{(n)})$, and $D_t^{(n)} = [1 - (1 + Y_{c,t}^{(n)})^{-n}]/[1 - (1 + Y_{c,t}^{(n)})^{-1}]$ is the approximate duration on an n-year coupon bond.

the left-hand side. I then report the coefficient on DUR_t from this specification as well as the associated t-statistic, labeled as "Difference from USA" in the table. Panel A shows results for DUR_t^{MBS}, and Panel B shows results for $DUR\_CNTRB_t^{MBS}$.

Table 7 shows that evidence exists of moderate cross-country spillovers. US MBS duration has some limited ability to forecast excess returns on long-term government bonds in Canada, Denmark, Japan, and the UK. However, in each case, the magnitude of the forecasting relations for foreign bonds is much weaker, both economically and statistically, than the corresponding magnitude for US bonds. In short, the evidence is consistent with the view that the bond markets of, say, the US and the UK are partially integrated, so a duration supply shock in the US has some impact on bond risk premia in the UK. At the same time, the fact that the US return forecasting results are so much stronger than the foreign return forecasting results provides additional comfort that the results do not reflect an omitted global risk factor.

3.4. Excess sensitivity of long-term yields and forwards to short rates

The model suggests that distant forward rates should be excessively sensitive to movements in short-term interest rates (see Proposition 2). As shown in Hanson and Stein (2012), this general excess sensitivity result emerges very strongly in the data. This is consistent with Gürkaynak, Sack, and Swanson (2005), who find that distant forwards are excessively sensitive to macroeconomic announcements. And, as emphasized by Hanson and Stein (2012), one observes significant excess sensitivity for distant real forward rates, i.e., forwards extracted from the Treasury Inflation Protected Securities (TIPS) yield curve. While this is puzzling from the standpoint of textbook macro-finance theories, excess sensitivity follows naturally from models that emphasize supply and demand effects in bonds markets. For instance, an extension in MBS duration increases the exposure of investors to movements in short-term real rates, necessitating a rise in the real term premium.

Going further, the model predicts that distant forwards should be particularly sensitive to short rates when the MBS market is more negatively convex (see Proposition 3). When the market is more negatively convex, a given change in short rates has a larger effect on the total quantity of interest rate risk and, hence, a larger impact on distant forward term premia. To test this prediction, I examine whether the high-frequency sensitivity of distant forwards varies with MBS convexity. In Table 8, I estimate daily regressions of the form

$$\Delta f_t^{X(10)} = a + b \cdot \Delta y_t^{\$(2)} + c \cdot NCONV_t^{MBS} + d \cdot (\Delta y_t^{\$(2)} \times NCONV_t^{MBS}) + \Delta \varepsilon_t^{X(10)}, \qquad (25)$$

where $NCONV_t^{MBS} = -CONV_t^{MBS}$ (negative one times the convexity of the Barclays MBS Index) for $X = \$$ and $TIPS$; i.e., I examine changes in both ten-year nominal and ten-year real forwards. I estimate these regressions from 1999 to the present to include data on real forwards from Gürkaynak, Sack, and Wright (2010).[23] The theory predicts that $d > 0$. Excess sensitivity

[23] The convexity of Barclays MBS Index is available for 1997–present. I obtain similar results for nominal forwards for this period.

should be more pronounced when the MBS market is more negatively convex.

[Insert Table 8 about here]

Consistent with this prediction, Table 8 shows that both distant nominal and real forwards are more sensitive to movements in short-term nominal rates when the MBS market is more negatively convex. Because $NCONV_t^{MBS}$ is persistent at daily frequencies, t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation up to 20 business days. From 1999 to the present, $NCONV_t^{MBS}$ has averaged 1.528 with a standard deviation of 0.531. Thus, on average, a 100 bps increase in the two-year nominal yield has been associated with a 54 bps [$= 0.322 + 0.140 \times 1.528$] increase in ten-year nominal forwards and a 30 bps increase in ten-year real forwards. However, when $NCONV_t^{MBS}$ is 2 standard deviations above average, the same 100 bps increase in the nominal short rates has been associated with a 68 bps increase in ten-year nominal forwards and a 42 bps increase in ten-year real forwards. Thus, variation in MBS convexity has been associated with meaningful variation in the sensitivity of distant forwards to movements in short rates. Furthermore, this is true whether I look at $NCONV_t^{MBS}$ or the scaled version of this variable, $NCONV_t^{MBS} \cdot \left(MV_t^{MBS}/MV_t^{AGG}\right)$, or I examine sensitivity to movements in short-term nominal yields or short-term forwards [if I replace $\Delta y_t^{\$(2)}$ with $\Delta f_t^{\$(2)}$ in Eq. (25)].

3.5. Excess volatility of long-term interest rates

The model suggests that, all else being equal, long-term interest rate volatility is elevated when the MBS market is more negatively convex. Intuitively, the positive-feedback dynamic underlying MBS duration spirals is stronger when the mortgage market is on a refinancing cliff; i.e., when a small movement in rates significantly impacts aggregate refinancing behavior.

To test this prediction, I follow the approach of Perli and Sack (2003), who relate weekly observations of option-implied interest rate variance to measures of MBS convexity. Specifically, Perli and Sack (2003) assume that $\Delta y_t = \sqrt{1+\beta x_t} \cdot \varepsilon_t$, where $\sigma^2_{\varepsilon,t}$ follows an $AR(1)$ process $\sigma^2_{\varepsilon,t} = \alpha_0 + \alpha_1 \sigma^2_{\varepsilon,t-1} + u_t$, so that

$$\sigma^2_{\Delta y,t} = (1+\beta x_t)\,\sigma^2_{\varepsilon,t} = \alpha_0\,(1+\beta x_t) + \alpha_1 \frac{1+\beta x_t}{1+\beta x_{t-1}} \sigma^2_{\Delta y,t-1} + (1+\beta x_t)\,u_t. \tag{26}$$

There are three effects: a level effect, a persistence effect, and a volatility of volatility effect. However, if x_t is fairly persistent, then $(1+\beta x_t)\,/\,(1+\beta x_{t-1}) \approx 1$. Furthermore, Perli and Sack (2003) find that the volatility of volatility effect is ignorable in practice. Thus, using $x_t = NCONV_t$, I focus solely on the level effect

$$\sigma^2_{\Delta y,t} = \alpha_0\,(1+\beta \cdot NCONV_t) + \alpha_1 \sigma^2_{\Delta y,t-1} + u_t. \tag{27}$$

A simultaneity problem arises here as $NCONV_t$ is a function of $\sigma^2_{\Delta y,t}$. To deal with this, I substitute $NCONV_{t-1}$ for $NCONV_t$ in the above.

Following Perli and Sack (2003), I measure $\sigma^2_{\Delta y,t}$ using the implied yield variance (the square of implied volatility) from three-month by ten-year swaptions (i.e., the option to enter into a ten-year swap at prespecified rate in three months). I first regress the implied variance from three-month by ten-year swaptions on the implied variance from two-year by ten-year swaptions and examine the residuals from this regression. As Perli and Sack (2003) argue, it is desirable to strip out structural or cyclical fluctuations in interest rate volatility in this way because MBS convexity is expected to have fairly transient effects on implied volatility. Using the resulting residualized measure, $\widetilde{\sigma}^2_{\Delta y,t}$, I then estimate

$$\widetilde{\sigma}^2_{\Delta y,t-1} = \alpha_0 \left(1 + \beta \cdot NCONV_{t-1}\right) + \alpha_1 \widetilde{\sigma}^2_{\Delta y,t-1} + u_t. \tag{28}$$

Data on MBS convexity are available beginning in 1997, so I estimate specification (28) using weekly data from 1997 to the present. t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at a lag of up to 12 weeks.

The results are shown in Table 9 and indicate that option-implied yield variance is higher when the MBS market is more negatively convex. This conclusion holds whether I look at $NCONV_t^{MBS}$ or the scaled version of this variable, $NCONV_t^{MBS} \cdot \left(MV_t^{MBS}/MV_t^{AGG}\right)$, and when I control for lagged swap yields in Eq. (28). The estimates in Column 2 imply that a 1 standard deviation increase in $NCONV$ raises implied yield volatility by 0.10% [$=\sqrt{0.020 \cdot 0.519}$]. The standard deviation of $\widetilde{\sigma}_{\Delta y,t-1}$ is 0.48%, suggesting that movements in MBS convexity are associated with economically significant shifts in interest rate volatility.

[Insert Table 9 about here]

3.6. Impact of MBS duration on corporate and swap spreads

In practice, MBS investors hedge the interest rate exposure of MBS using either Treasuries or interest rate swaps. Regardless of the specific hedging instrument, other investors must take the other side of these trades. Thus, in the aggregate, fixed income investors must bear a time-varying amount of interest rate risk, which impacts equilibrium term premia. In a simple model in which Treasuries and swaps are perfect substitutes, one would not expect these hedging flows to impact spreads between swaps and Treasuries. But different long-term fixed income assets with the same duration are not perfect substitutes. Once we relax the assumption of perfect substitutability, hedging flows triggered by shifts in MBS duration could impact yield spreads between duration-matched fixed income assets and give rise to predictable variation in the excess returns on various fixed income assets over duration-matched government bonds.

In the Internet Appendix, I consider a stylized model with three long-term fixed income assets: government bonds, interest rate swaps, and high-grade corporate bonds. I suppose that some fraction of MBS duration is hedged using interest rate swaps and the remainder is hedged using government bonds. While all long-term fixed income assets are exposed to interest rate risk, some components of returns are specific to government bonds, swaps, and corporate bonds.

For corporate bonds, this could reflect changes in credit risk. And, for all three assets, these components could be reflecting shifts in supply and demand for specific assets. For instance, in the case of government bonds, this could reflect idiosyncratic shifts in the demand as in a flight to quality episode (e.g., Duffee, 1996; Longstaff, 2004; etc.) or idiosyncratic shifts in supply (e.g., Greenwood and Vayanos, 2010; and Lou, Yan, and Zhang, 2013).

In my stylized model, the expected excess returns on any long-term fixed income asset over short-term bonds is equal to a term premium earned by all long-term fixed income assets plus a risk premium that is specific to that asset. In particular, because long-term government bonds and swaps are both exposed to movements in the general level of interest rates, the impact of MBS duration on the overall term premium is independent of the fraction of MBS hedged with government bonds and swaps—just as in the simpler model developed above.

However, MBS hedging flows now have an impact on the spread between long-term fixed income assets. For instance, the government-specific risk premium is high when MBS duration is high. Because MBS investors hedge with government bonds, the government-specific risk premium must rise to induce arbitrageurs to hold more Treasuries. Because these hedging flows have no impact on the corporate-specific risk premium, they result in tighter than normal spreads between long-term corporate and government bonds. Furthermore, if a sufficiently large volume of MBS hedging takes place in swap markets, and the idiosyncratic movements in swaps are large relative to the idiosyncratic movements in Treasuries, an increase in MBS duration would be expected to raise the spread between swaps and Treasuries.

In summary, the analysis suggests that high levels of MBS duration are associated with narrow current spreads between corporate bonds and duration-matched Treasuries, should forecast a future widening of corporate bond spreads, and should predict low future returns on corporate bonds relative to those on duration-matched Treasuries. And because swaps are used to hedge MBS, each of these predictions should be reversed for swaps.

I find evidence consistent with each of these additional predictions in Table 10. In Panel A, I report regressions of the form

$$\Delta SPREAD_t = a_1 + b_1 \cdot \Delta DUR\_CNTRB_t^{MBS} + \mathbf{c}_1' \Delta \mathbf{x}_t + v_t. \tag{29}$$

Specifically, I regress three-month changes in spreads on the contemporaneous change in MBS duration contribution and controls. I find that increases in MBS duration are associated with tighter spreads between corporate bonds and Treasuries and with wider swap spreads. This holds in univariate regressions as well as in multivariate regressions that control for changes in the term spread, the Chicago Board Options Exchange Market Volatility Index (VIX), and the past returns on the stock market, which Collin-Dufresne, Goldstein, and Martin (2001) show are useful for explaining changes in corporate credit spreads. Spreads are in percentage points, so the coefficient of -0.43 in Column 1 suggests that a one-year rise in $DUR\_CNTRB_t^{MBS}$ is associated with a 43 bps decline in investment-grade spreads. Based on the univariate estimates, a 1 standard deviation change in $DUR\_CNTRB_t^{MBS}$ is associated with a 11 bps decline in all investment-grade corporate spreads, a 4 bps decline in Aaa spreads, a 19 bps decline in Baa spreads, and a 3 bps rise in ten-year swap spreads.

[Insert Table 10 about here]

In Panel B, I use the level of MBS duration to forecast spread changes over the next 12 months:

$$\Delta SPREAD_{t+1} = a_2 + b_2 \cdot DUR\_CNTRB_t^{MBS} + \mathbf{c}_2'\mathbf{x}_t + u_{t+1}. \tag{30}$$

I find that a high level of MBS duration predicts that corporate spreads will widen over the next 12 months and that swap spread spreads will tighten. Again, this holds true in univariate forecasting regressions as well as in multivariate regressions that control for the initial level of spreads and other conditioning variables.

Finally, I forecast the excess returns on corporate bonds and swaps over duration-matched Treasuries over the following 12 months:

$$rx_{t+1}^{DUR-MATCH} = a_3 + b_3 \cdot DUR\_CNTRB_t^{MBS} + \mathbf{c}_3'\mathbf{x}_t + \varepsilon_{t+1}. \tag{31}$$

By construction, these regressions provide almost identical information to those in Panel B. Specifically, a high level of MBS duration predicts that the returns on long-term corporate bonds will underperform those on duration-matched Treasuries, while swaps are expected to outperform duration-matched Treasuries.

Finally, Perli and Sack (2003), Wooldridge (2001), and Reinhart and Sack (2000) argue that many MBS investors switched from hedging with Treasuries to hedging with interest rate swaps beginning in the late 1990s. Several episodes, including the flight-to-quality dislocations in the fall of 1998 and the jump in the convenience premium on Treasuries following the Treasury buyback announcement in early 2000, convinced many MBS investors that hedging with Treasuries exposed them to far greater basis risk than hedging with swaps. Interestingly, the time series evidence is consistent with such a shift in hedging activity. In untabulated results, I find that the tendency for swap spreads to widen when MBS duration rises became significantly more pronounced after 1997.

In summary, the evidence is consistent with previous work, including Cortes (2003, 2006) and Feldhutter and Lando (2008), which argues that MBS hedging plays a significant role in explaining the level of US swap spreads. And this spread-based evidence supports the broader argument of this paper: Shifts in MBS duration and the associated hedging flows function as large-scale supply shocks that have significant effects on bond market pricing.

4. Conclusion

Changes in the effective duration of mortgage-backed securities function as large-scale duration supply shocks in US bond markets. As a result, bond term premia are high when aggregate MBS duration is high, and changes in MBS duration capture variation in bond risk premia that is not reflected in traditional forecasting variables. The fact that these recurring and transient supply shocks have large effects on bond prices highlights the critical role of limited and slow-moving arbitrage capital, even in markets as deep and liquid as the US bond market.

The negative convexity of MBS—the fact that MBS duration rises when interest rates rise—generates a positive-feedback loop that helps explain the excess sensitivity of long rates to short rates and excess volatility of long rates. MBS convexity has the potential to amplify a variety of shocks within US bond markets. For instance, one might expect the effective risk tolerance of fixed income arbitrageurs to decline following losses. If arbitrageurs are long duration and lose money when rates rise, negative MBS convexity could add another positive-feedback loop to US bond markets. Specifically, a decline in arbitrageur risk tolerance would lead to a rise in yields, which would cause the duration of MBS to extend, raising required returns and causing yields to rise further. The resulting losses would further lower arbitrageur risk tolerance, and so on.

Thus, MBS convexity could sometimes act as a destabilizing force within US bond markets, as it arguably did at the beginning of the three most recent Fed tightening cycles in 1994, 1999, and 2003. While it could be desirable for households to have access to mortgages that embed a no-penalty prepayment option, this could come at some cost in terms of financial stability. More speculatively, my analysis offers a potential explanation for the finding that failures of the expectations hypothesis are more pronounced in the US than in most other developed nations (see, e.g., Campbell, 2003). Because the US is one of the few countries with long-term fixed rate mortgages that embed a true interest rate call, the existence of a significant MBS convexity effect suggests a novel solution to this puzzle.

Finally, my analysis indicates that shocks to MBS duration impact term premia throughout the US bond market (for government debt, MBS, and corporate debt alike) as predicted by a model such as Vayanos and Vila (2009) in the portfolio-balance tradition. This finding is relevant for the ongoing debate about the effectiveness of Quantiative Easing policies, the purchases of long-term bonds by the Federal Reserve and other global central banks. For instance, Krishnamurthy and Vissing-Jorgensen (2013) argue that the markets for MBS, corporate bonds, and Treasuries are highly segmented from one another, so QE policies do not work through broad channels such a shifting the market-wide price of interest rate risk. My results do not show that QE policies have lowered the duration risk premium in bond markets. However, they do provide something of a proof of concept for QE policies that seek to impact market-wide term premia. Comparably sized shocks to the quantity of duration appear to have had significant effects on term premia in the past.

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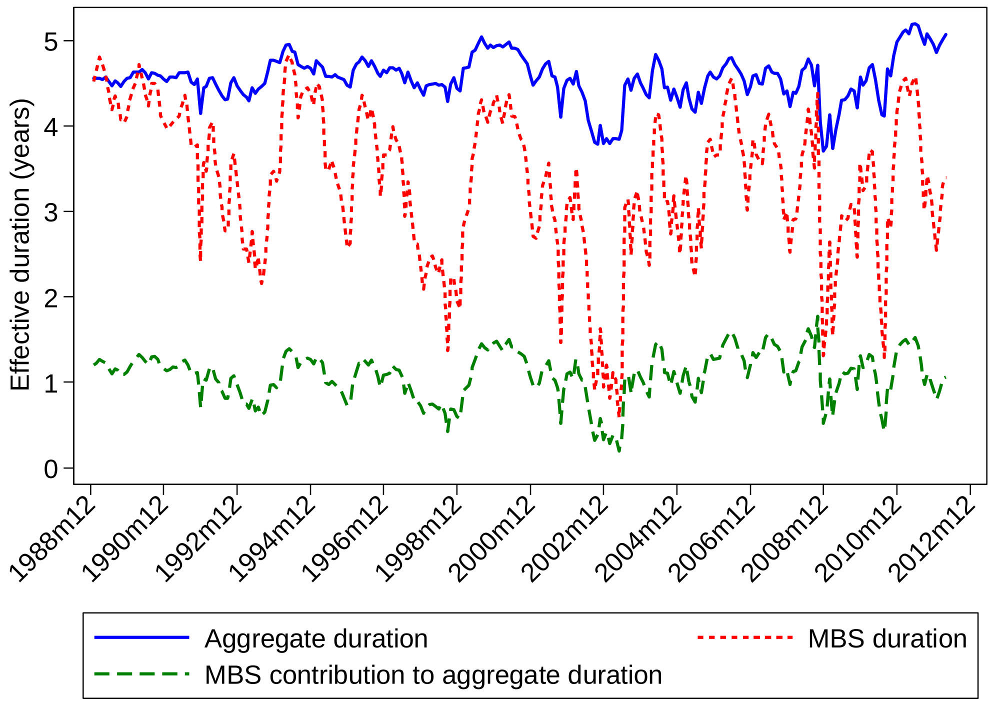
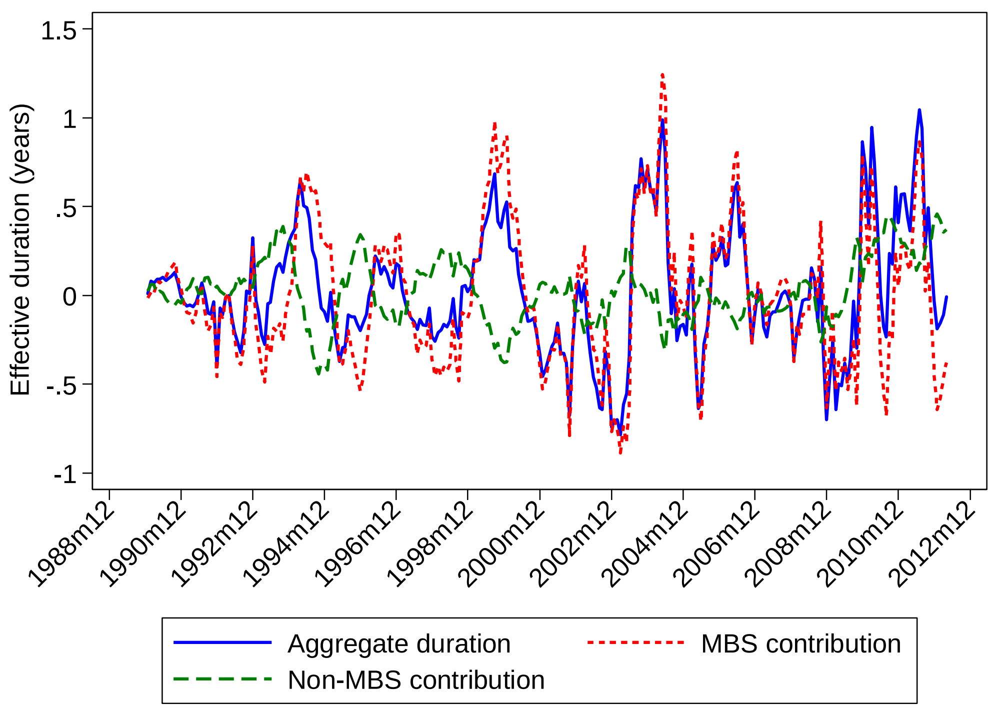
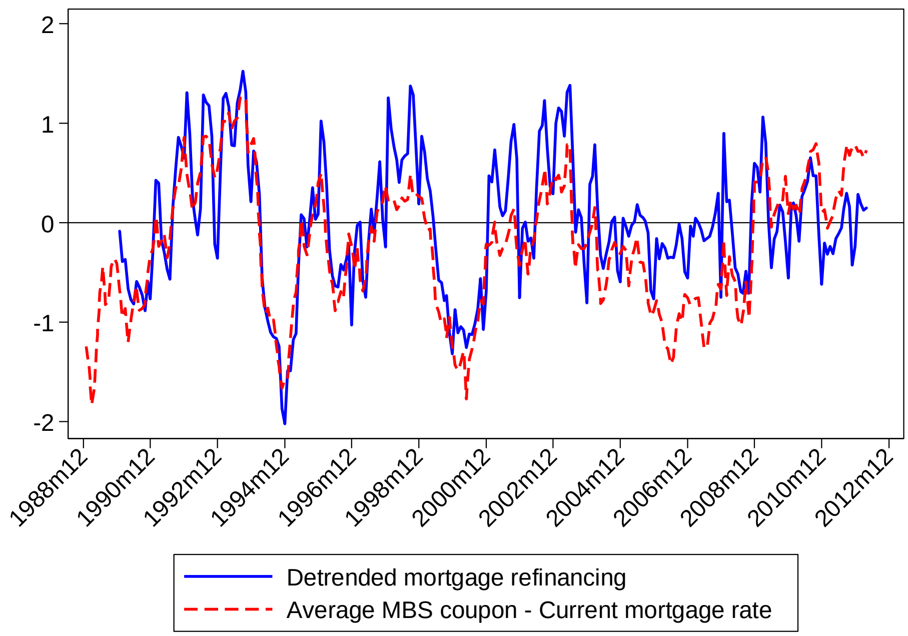
Panel A: Effective duration in levels



Panel B: 12-month changes in effective duration



Panel C: Detrended refinancing activity and refinancing incentives



Panel D: Detrended duration in ten-year US Treasury equivalents (in billions of 2012 dollars)

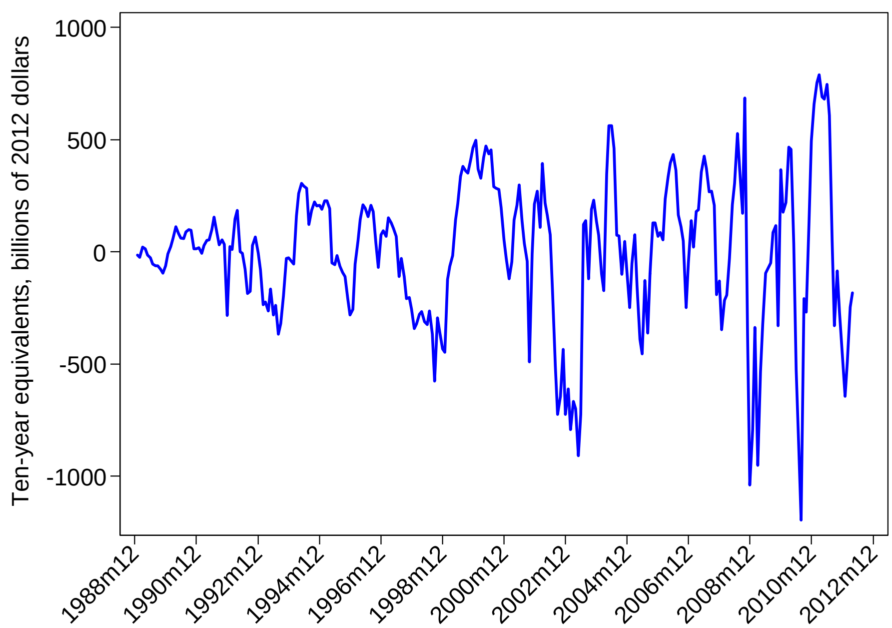
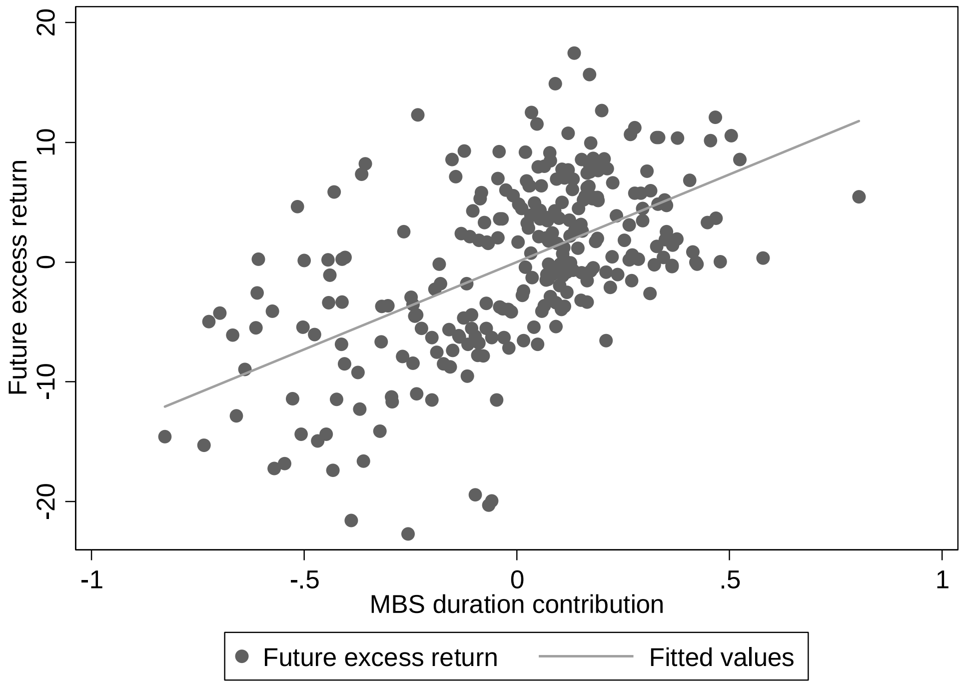


Fig. 1. Effective duration of the US fixed income market. Panel A plots effective duration measures based on Barclays bond indices from 1989m1–2012m4: the effective duration of the Barclays Aggregate Index (DUR^{AGG}), the effective duration of the Barclays Mortgage-backed Securities (MBS) Index (DUR^{MBS}), and the duration contribution of the MBS Index to Aggregate duration (DUR_CNTRB^{MBS}). Panel B shows the decomposition from Eq. (11) in 12-month changes. Specifically, the change in Aggregate duration is the sum of the MBS contribution and the non-MBS contribution. Panel C plots detrended refinancing activity based on the Mortgage Bankers' Association Refinancing Index versus aggregate refinancing incentives (the average coupon on outstanding MBS minus the primary mortgage rate). Panels D shows detrended MBS duration in ten-year Treasury equivalents (in billions of 2012 dollars).

Panel A: Scatter plot



Panel B: Time-series plot

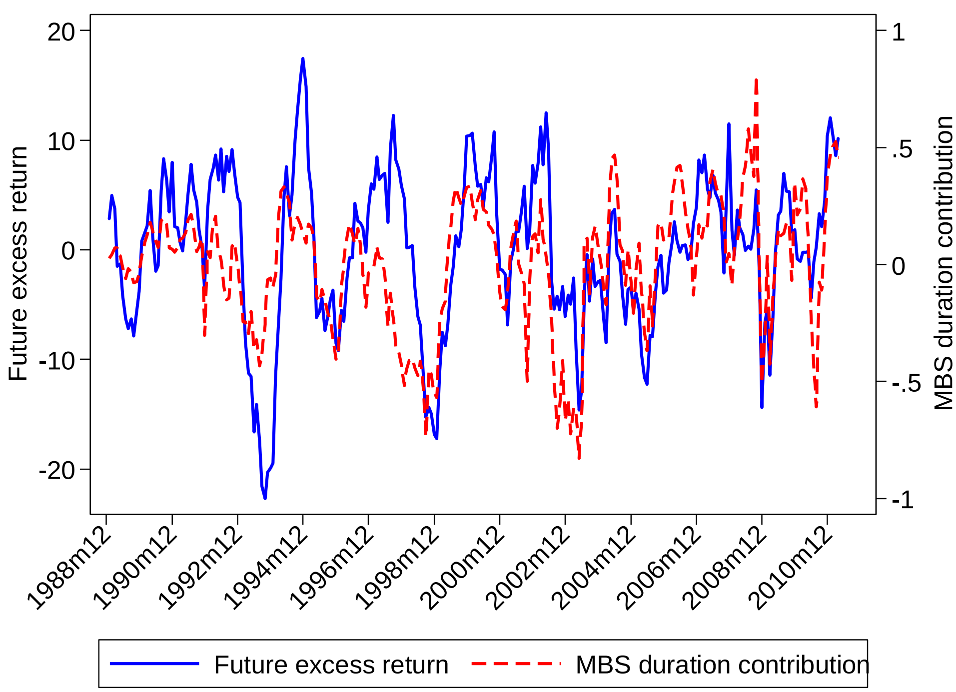
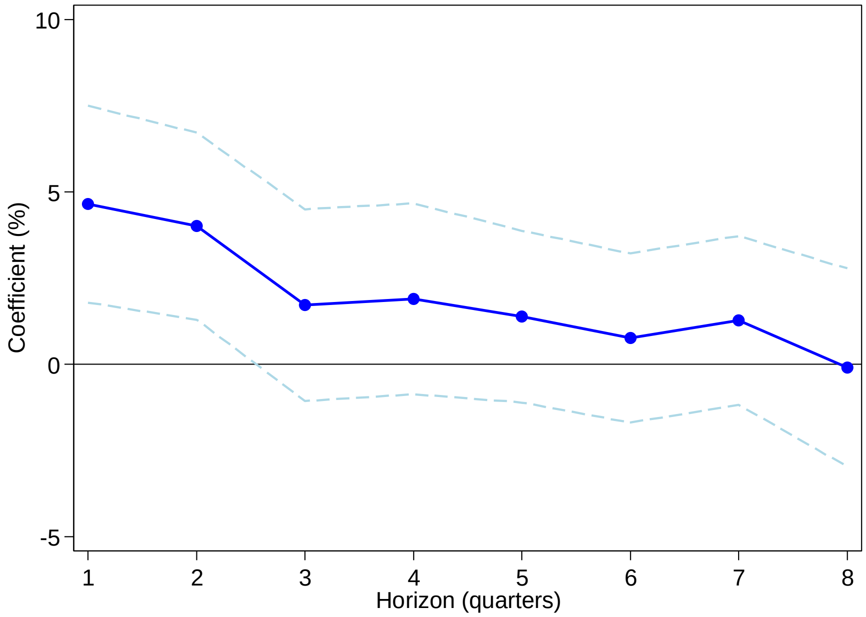


Fig. 2. The basic forecasting result using mortgage-backed securities (MBS) duration contribution. This figure depicts the regression of future 12-month excess bond returns on the term spread and $DUR\_CNTRB_t^{MBS}$, corresponding to Column 10 of Table 2. Specifically, Panel A shows a scatter plot of the component of future excess returns that is orthogonal to the term spread versus the component of $DUR\_CNTRB_t^{MBS}$ that is orthogonal to the term spread. Panel B shows the corresponding time series plot.

Panel A: Univariate coefficients by quarterly forecast horizon



Panel B: Multivariate coefficients by quarterly forecast horizon

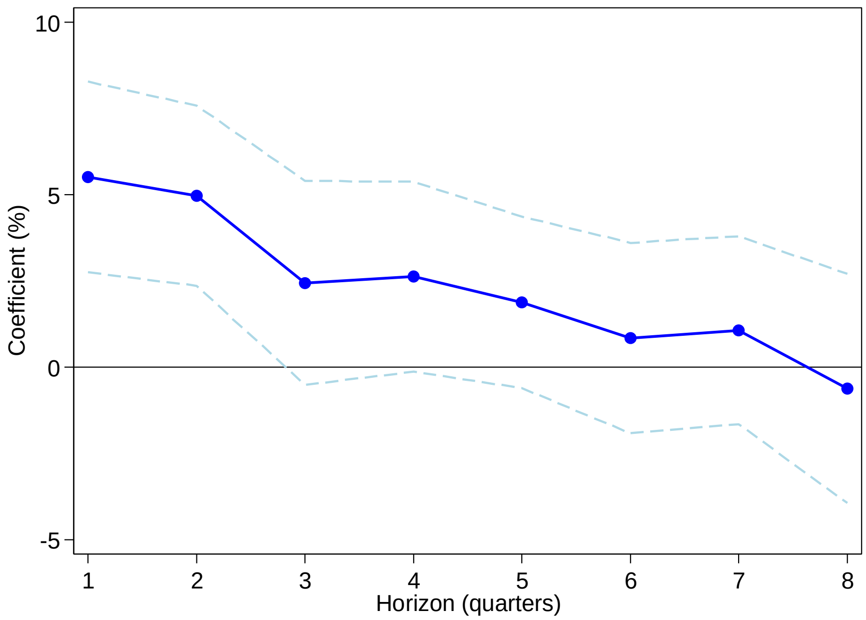


Fig. 3. Coefficients by quarterly forecast horizons. This figure plots the coefficients $b_{(j)}$ on MBS duration contribution from estimating

$$rx^{(10)}_{t+(j-1)/4\rightarrow t+j/4} = a_{(j)} + b_{(j)} \cdot DUR\_CNTRB^{MBS}_{t} + \mathbf{c}'_{(j)}\mathbf{x}_t + \varepsilon^{(10)}_{t+(j-1)/4\rightarrow t+j/4}$$

for quarters $j = 1, ..., 8$, effectively tracing out a nonparametric version of the impulse to quarterly excess bond returns following a movement in mortgage-backed securities (MBS) duration. Panel A plots the coefficients from univariate forecasting regressions. The multivariate specifications in Panel B controls for the term spread. Confidence intervals, based on Newey and West (1987) standard errors allowing for serial correlation at up to six months, are shown as dashed lines.

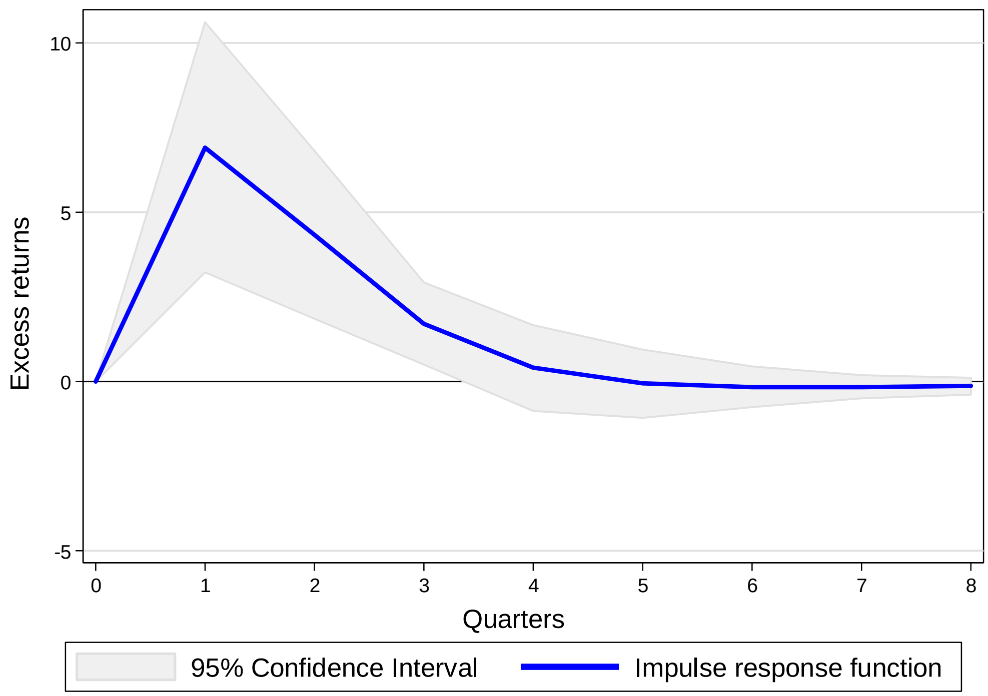
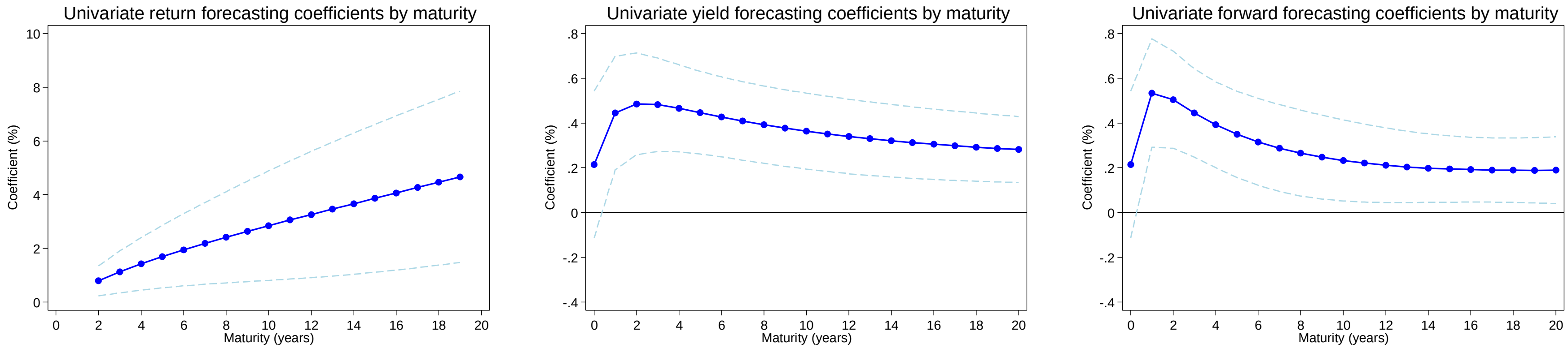


Fig. 4. Impulse response of quarterly excess bond returns to shock to mortgage-backed securities (MBS) duration. Using quarterly data, I estimate a first order vector autoregression (VAR) of the form

$$\begin{aligned} rx_{t+1/4}^{(10)} &= a_1 + b_1 \cdot rx_t + c_1 \cdot DUR_t + d_1 \cdot (y_t^{(10)} - y_t^{(1)}) + \varepsilon_{1,t+1/4}^{(10)} \\ DUR_{t+1/4} &= a_2 + b_2 \cdot rx_t + c_2 \cdot DUR_t + d_2 \cdot (y_t^{(10)} - y_t^{(1)}) + \varepsilon_{2,t+1/4}^{(10)} \\ y_{t+1/4}^{(10)} - y_{t+1/4}^{(1)} &= a_3 + b_3 \cdot rx_t + c_3 \cdot DUR_t + d_3 \cdot (y_t^{(10)} - y_t^{(1)}) + \varepsilon_{3,t+1/4}^{(10)} \end{aligned}$$

For simplicity, I show results only for $DUR = DUR\_CNTRB^{MBS}$. I plot the (simple) impulse response function from a shock to DUR_CNTRB^{MBS} in quarter $t = 0$ on excess bond returns from quarters $t = 1$, ..., 8. Confidence intervals for the estimated impulse response are shown in gray. For simplicity, I use a quarterly data set of nonoverlapping observations, sampled in March, June, September, and December.

Panel A: MBS duration



Panel B: MBS duration contribution

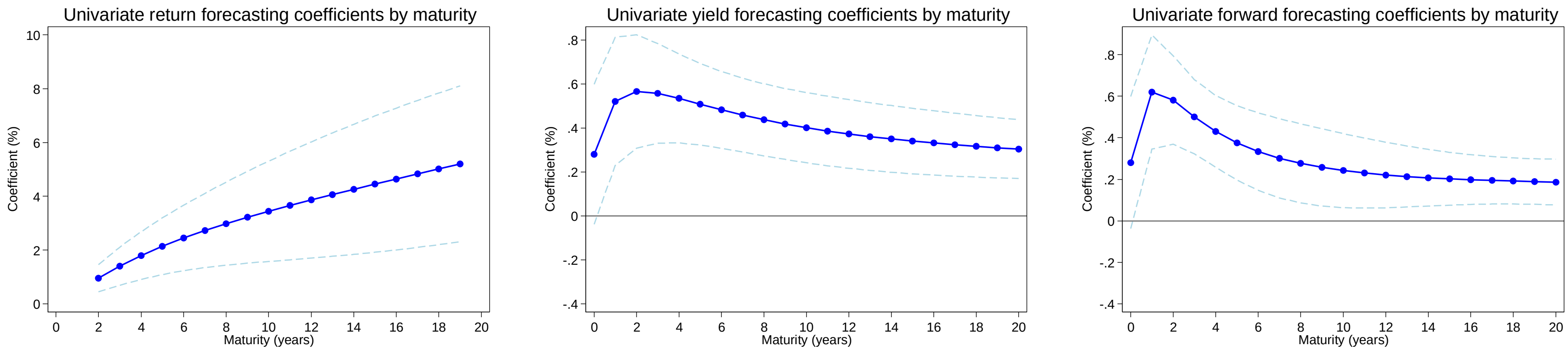


Fig. 5. Excess return, yield, and forward rate forecasting coefficients by bond maturity. This figure plots the coefficients $b_{(n)}$ from estimating

$$rx_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)}$$

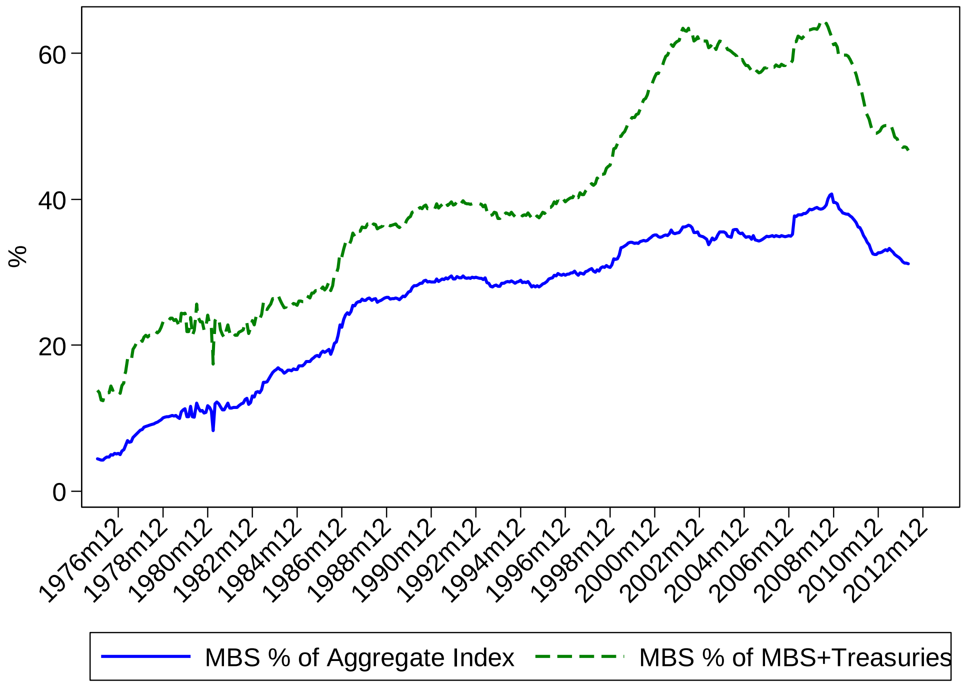
versus maturity n, for $n = 2,\ldots,20$. I then plot the coefficients $b_{(n)}$ versus maturity n from estimating

$$y_t^{(n)} - y_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)}$$

$$f_t^{(n)} - f_{t+1}^{(n)} = a_{(n)} + b_{(n)} \cdot DUR_t + \varepsilon_{t+1}^{(n)},$$

for $n = 0, 1, 2,\ldots,20$. For simplicy, I rescale the coefficients $b_{(n)}$ so that they reflect the impact of one standard deviation shift in DUR_t. Confidence intervals, based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 months, are shown as dashed lines. Panel A shows results for mortgage-backed securities (MBS) duration, and Panel B shows results for MBS duration contribution. In each panel, the left-most graph shows the excess return forecasting results by maturity, the middle graph shows yield-change forecasting results by maturity, and the right-most graph shows forward-change forecasting results.

Panel A: The growth of the MBS market



Panel B: MBS duration measures

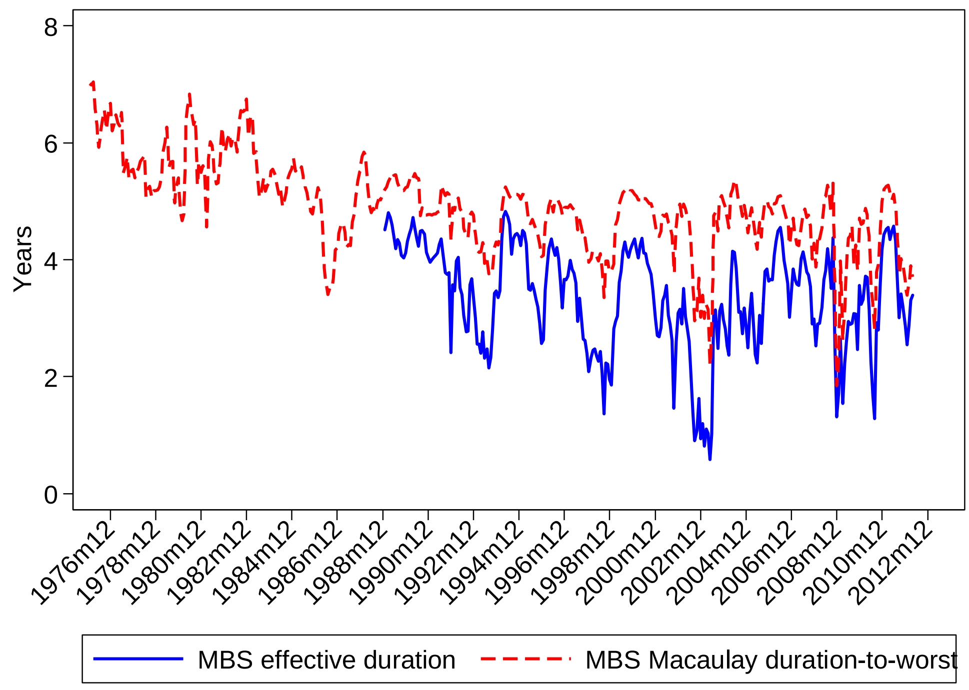
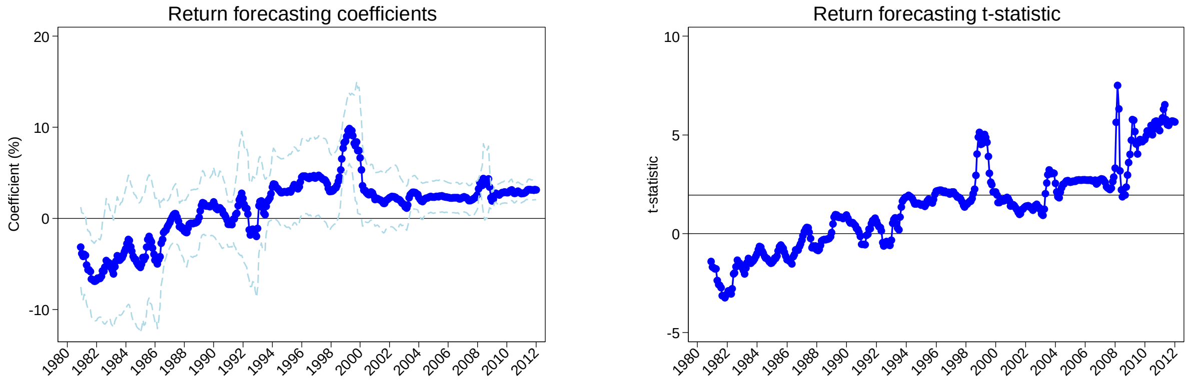
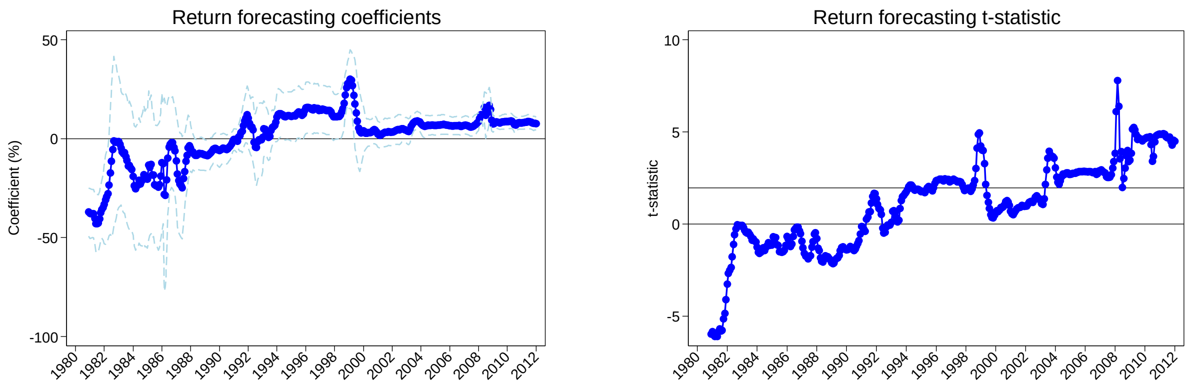


Fig. 6. Growth of mortgage-backed securities (MBS) market and MBS duration measures. Panel A shows the market value of the Barclays MBS Index as a fraction of the market value of the Barclays Aggregate Index (a proxy for broad the investment-grade bond market) from 1976–2012 and as a fraction of the sum of the value of the Barclays MBS and Treasury indices. Panel B compares the effective duration of the Barclays MBS Index from 1989–2012 to the Macaulay duration-to-worst of the index from 1976–2012.

Panel A: MBS Macaulay duration ($dur = dur^{MBS}$)



Panel B: MBS Macaulay duration contribution ($dur = dur\_cntrb^{MB}$)



Panel C: Aggregate refinancing incentives ($dur = c - y_M$)

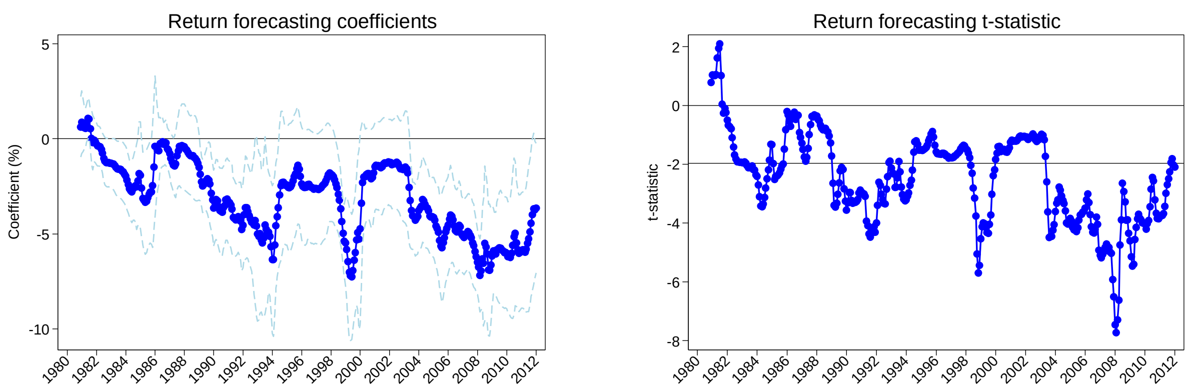


Fig. 7. Coefficients from 60-month rolling excess return forecasting regressions, 1980m12-2012m1. This figure depicts results from 60-month rolling regressions of three-month excess bond returns on mortgage-backed securities (MBS) duration measures, controlling for the term spread

$$rx^{(10)}_{t+1/4} = a + b \cdot dur_t + c \cdot (y^{(10)}_t - y^{(1)}_t) + \varepsilon^{(10)}_{t+1/4}.$$

Panel A shows results for MBS Macaulay duration ($dur = dur^{MBS}$), Panel B shows results for MBS Macaulay duration contribution ($dur = dur\_cntrb^{MB}$), and Panel C shows results for aggregate refinancing incentives ($dur = c - y_M$). The left-hand graph in each Panel shows the estimated regression coefficients on *dur* along with confidence intervals as dashed lines. The right-hand graph shows the associated *t*-statistics. I compute standard errors using Newey and West (1987) standard errors allowing for serial correlation at up to six months.

Table 1
Summary statistics.
This table presents means, medians, standard deviations, extreme values, and monthly time series autocorrelations (denoted ρ) of variables between 1989m1 and 2011m4. Panel A presents summary statistics for US Treasury excess returns, yield spreads, and forwards from Gürkaynak, Sack, and Wright (2007). These variables are all measured in percentage points. Panel B presents summary statistics for various measures of effective duration in years based on Barclays Capital bond indices. Panel C reports the corresponding summary statistics for effective convexity. The effective convexity measures are available beginning only in 1997m4.

Variable	N	Mean	Median	St Dev	Min	Max	ρ
Panel A: US Treasury excess returns, yield spreads, and forwards (%)							
$rx_{t+1}^{(5)}$	268	3.00	3.33	4.39	-8.49	12.46	0.922
$rx_{t+1}^{(10)}$	268	4.98	5.74	7.73	-16.41	20.96	0.884
$rx_{t+1}^{(20)}$	268	6.32	6.35	12.98	-29.35	39.17	0.861
$y_t^{(5)} - y_t^{(1)}$	268	0.90	0.76	0.83	-0.54	2.50	0.973
$y_t^{(10)} - y_t^{(1)}$	268	1.57	1.20	1.28	-0.58	3.79	0.977
$y_t^{(20)} - y_t^{(1)}$	268	2.01	1.57	1.54	-0.82	4.68	0.978
$f_t^{(1)}$	268	4.45	4.59	2.24	0.16	9.70	0.976
$f_t^{(2)}$	268	4.93	4.98	2.01	0.62	9.32	0.971
$f_t^{(3)}$	268	5.33	5.35	1.80	1.42	9.19	0.970
$f_t^{(4)}$	268	5.67	5.60	1.64	2.31	9.17	0.970
$f_t^{(5)}$	268	5.96	5.76	1.52	3.02	9.16	0.969
Panel B: Effective duration (years)							
DUR_t^{AGG}	268	4.53	4.56	0.26	3.71	5.12	0.863
DUR_t^{MBS}	268	3.35	3.49	0.91	0.58	4.83	0.875
$DUR\_CNTRB_t^{MBS}$	268	1.08	1.12	0.29	0.20	1.78	0.851
DUR_t^{OTH}	268	5.13	5.16	0.25	4.47	5.61	0.954
$DUR\_CNTRB_t^{OTH}$	268	3.46	3.45	0.18	2.93	3.88	0.960
MV_t^{MBS}/MV_t^{AGG}	268	0.33	0.33	0.04	0.26	0.41	0.991
Panel C: Effective convexity (years)							
$CONV_t^{AGG}$	169	-0.18	-0.19	0.17	-0.52	0.20	0.844
$CONV_t^{MBS}$	169	-1.50	-1.54	0.45	-2.43	-0.44	0.825
$CONV_t^{MBS} \cdot (MV_t^{MBS}/MV_t^{AGG})$	169	-0.52	-0.54	0.16	-0.85	-0.15	0.818

Table 2

Forecasting excess bond returns using measures of mortgage-backed securities (MBS) duration.

Regressions of 12-month excess returns on ten-year Treasuries on the effective duration of the Barclays Aggregate Index, the effective duration of the Barclays MBS Index, and the effective duration contribution of the Barclays MBS Index:

$$rx_{t+1}^{(10)} = a + b \cdot DUR_t + \mathbf{c}'\mathbf{x}_t + \varepsilon_{t+1}^{(10)}.$$

The regressions are estimated with monthly data from 1989m1-2011m4, so each month I forecast the excess return over the following 12 months. To deal with the overlapping nature of returns, *t*-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. I estimate these regressions with and without other forecasting variables identified in the literature on bond risk premia. Specifically, I control for the term spread following Campbell and Shiller (1991) and the first five forward rates following Cochrane and Piazzesi (2005).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
DUR_t^{AGG}			6.725	11.482	7.749						
			[2.84]	[5.44]	[2.73]						
DUR_t^{MBS}						2.905	4.437	3.929			
						[2.76]	[6.48]	[4.59]			
$DUR\_CNTRB_t^{MBS}$									11.253	14.629	12.143
									[3.70]	[5.78]	[4.44]
$y_t^{(10)} - y_t^{(1)}$	2.178			2.928			3.231			2.999	
	[3.02]			[3.88]			[4.67]			[4.57]	
$f_t^{(1)}$		-37.854			-30.755			-27.953			-22.653
		[-6.77]			[-4.91]			[-4.59]			[-3.89]
$f_t^{(2)}$		217.606			171.892			146.201			114.797
		[7.00]			[4.77]			[4.68]			[3.78]
$f_t^{(3)}$		-511.359			-403.066			-340.160			-268.889
		[-6.92]			[-4.80]			[-4.74]			[-3.78]
$f_t^{(4)}$		533.181			418.391			357.217			286.175
		[6.66]			[4.72]			[4.60]			[3.65]
$f_t^{(5)}$		-201.349			-156.236			-136.066			-109.305
		[-6.26]			[-4.47]			[-4.29]			[-3.36]
Constant	1.555	2.065	-25.513	-51.686	-34.283	-4.740	-14.948	-6.634	-7.158	-15.515	-11.768
	[1.04]	[0.48]	[-2.34]	[-5.08]	[-2.56]	[-1.18]	[-4.58]	[-1.27]	[-1.85]	[-4.23]	[-1.95]
Number of observations	268	268	268	268	268	268	268	268	268	268	268
R^2	0.13	0.31	0.05	0.26	0.35	0.12	0.37	0.42	0.17	0.40	0.46

Table 3

Forecasting excess bond returns, mortgage-backed securities (MBS) vs. non-MBS duration.

Regressions of 12-month excess returns on ten-year Treasuries on the effective duration of the Barclays Aggregate Index, the effective duration contribution of the Barclays MBS Index, and the effective duration contribution of the non-MBS index:

$$rx_{t+1}^{(10)} = a + b \cdot DUR_t + \mathbf{c}'\mathbf{x}_t + \varepsilon_{t+1}^{(10)}.$$

The regressions are estimated with monthly data from 1989m1-2011m4, so each month I forecast the excess return over the following 12 months. To deal with the overlapping nature of returns, *t*-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. I estimate these regressions with and without other forecasting variables identified in the literature on bond risk premia. Specifically, I control for the term spread following Campbell and Shiller (1991) and the first five forward rates following Cochrane and Piazzesi (2005).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
DUR_t^{AGG}	6.725				11.482				7.749			
	[2.84]				[5.44]				[2.73]			
$DUR\_CNTRB_t^{MBS}$		11.253		8.872		14.629		13.586		12.143		10.178
		[3.70]		[3.25]		[5.78]		[6.63]		[4.44]		[3.48]
$DUR\_CNTRB_t^{OTH}$			-14.620	-8.022			-13.676	-3.179			-13.200	-5.577
			[-2.10]	[-1.11]			[-1.92]	[-0.44]			[-1.73]	[-0.68]
$y_t^{(10)} - y_t^{(1)}$					2.928	2.999	2.053	2.912				
					[3.88]	[4.57]	[2.74]	[4.14]				
$f_t^{(1)}$									-30.755	-22.653	-33.422	-23.239
									[-4.91]	[-3.89]	[-5.39]	[-4.00]
$f_t^{(2)}$									171.892	114.797	183.721	117.115
									[4.77]	[3.78]	[5.04]	[3.86]
$f_t^{(3)}$									-403.066	-268.889	-432.257	-274.699
									[-4.80]	[-3.78]	[-4.87]	[-3.96]
$f_t^{(4)}$									418.391	286.175	460.215	295.311
									[4.72]	[3.65]	[4.87]	[3.99]
$f_t^{(5)}$									-156.236	-109.305	-178.142	-114.393
									[-4.47]	[-3.36]	[-4.85]	[-3.82]
Constant	-25.513	-7.158	55.500	23.133	-51.686	-15.515	49.011	-3.268	-34.283	-11.768	48.946	10.278
	[-2.34]	[-1.85]	[2.31]	[0.88]	[-5.08]	[-4.23]	[1.98]	[-0.12]	[-.56]	[-1.95]	[1.92]	[0.34]
Number of observations	268	268	268	268	268	268	268	268	268	268	268	268
R^2	0.05	0.17	0.12	0.20	0.26	0.40	0.23	0.41	0.35	0.46	0.39	0.47

Table 4

Comparison with literature on the Federal Reserve's Large-Scale Asset Purchase (LSAP) programs.

This table compares the implied price impact from my regressions using mortgage-backed securities (MBS) duration with that estimated by papers examining the Fed's LSAP programs. I obtain the implied effect for each LSAP as

$$\text{Implied Effect in basis points (bps)} = (\text{Effect of 1 } \sigma \text{ move in } DUR_t \text{ on } y^{(10)} \text{ in bps}) \times \frac{\text{LSAP size in ten} - \text{year Treasury duration equivalents}}{\text{ten} - \text{year Treasury duration equivalents corresponding to 1 } \sigma \text{ move in } DUR_t}$$

Li and Wei (2013) estimate that the $300 billion of Treasury purchases associated with LSAP1 reduced duration supply by $169 billion = $(4.5 \times 300)/8$ billion ten-year Treasury equivalents. Thus, accounting for the purchases of $200 billion of agency debt and $1,250 billion MBS, I estimate that LSAP1 reduced the aggregate supply of duration by $750 billion = $(4.5 \times 300 + 4.5 \times 200 + 3 \times 1250)/8$ billion tem-year Treasury equivalents. Gagnon, Raskin, Remache, and Sack (2011) examine movements in yields on eight LSAP1 announcement dates and estimate that LSAP1 reduced ten-year yields by 91 bps. Applying the coefficients from time-series regressions, Gagnon, Raskin, Remache, and Sack (2011) estimate an impact of LSAP1 of 61 bps. Krishnamurthy and Vissing-Jorgensen (2011) examine movement in yields on five LSAP1 announcement dates and estimate that LSAP1 reduced ten-year yields by 107 bps. Using a no arbitrage model, Li and Wei (2013) estimate that LSAP1 lowered yields by 99 bps. Jarrow and Li (2012) also estimate a no arbitrage model and estimate that LSAP1 and LSAP2 lowered ten-year yields by 70 bps. Assuming that roughly 80% of this corresponds to LSAP1, this implies that LSAP1 lowered yields by 56 bps.

Li and Wei (2013) estimate that LSAP2 reduced the supply of duration by $400 billion ten-year Treasury equivalents. Krishnamurthy and Vissing-Jorgensen (2011) examine movement in yields on three LSAP2 announcement dates and estimate that LSAP2 reduced ten-year yields by 30 bps. Using a no-arbitrage model, Li and Wei (2013) estimate that LSAP2 lowered yields by 26 bps. Jarrow and Li (2012) also estimate a no-arbitrage model and estimate that LSAP1 and LSAP2 lowered ten-year yields by 70 bps. Assuming roughly 20% of Jarrow and Li's (2012) estimate corresponds to LSAP2, this implies that LSAP2 lowered yields by 14 bps.

LSAP3 is often referred to as the Maturity Extension Program. Li and Wei (2013) estimate that LSAP3 reduced the supply of duration by $400 billion ten-year Treasury equivalents. Li and Wei (2013) estimate that LSAP3 lowered ten-year yields by 25 bps. Using regressions similar to those in Greenwood and Vayanos (2013), Hamilton and Wu (2012) estimate that LSAP3 lowered ten-year yields by 17 bps.

		Supply shock	Estimated effect on ten-year yield (bps)						
			LSAP evaluation studies			Effect implied by MBS supply shocks			
	Event description	Ten-year Treasury duration equivalents	Average study	Minimum study	Maximum study	DUR_t^{MBS} Univariate	DUR_t^{MBS} Multivariate	$DUR\_CNTRB_t^{MBS}$ Univariate	$DUR\_CNTRB_t^{MBS}$ Multivariate
LSAP1	Purchase of $300billion Treasuries, $200 billion Agencies, $1,250 billion MBS. Announced 11/2008–3/2009.	$750 billion	83	55	107	54	68	69	79
LSAP2	Purchase of $600 billion of long-term Treasuries. Announced 11/2010.	$400 billion	24	15	30	29	36	37	42
LSAP3	Buy $400 billion long-term Treasuries and sell $400 billion short-term Treasuries. Announced 9/2011.	$400 billion	21	17	25	29	36	37	42
Effect of 1 σ move in DUR_t on $y^{(10)}$ in bps						36	46	40	46
Ten − year Treasury duration equivalents corresponding to 1 σ move in DUR_t						503	503	436	436

Table 5
Horse races between mortgage-backed securities (MBS) duration and mortgage refinancing measures.
Regressions of 12-month excess returns on ten-year Treasuries on the effective duration contribution of the Barclays MBS Index and either aggregate refinancing incentives ($c_t - y_{M,t}$) computed as the difference between the average MBS coupon and the Freddie Mac primary market 30-year mortgage rate or the log-deviation of the Mortgage Bankers' Association refinancing index from its Hodrick-Prescott (1997) trend ($REFI_t$):

$$rx_{t+1}^{(10)} = a + b \cdot DUR\_CNTRB_t^{MBS} + c \cdot REFI_t + \mathbf{d}'\mathbf{x}_t + \varepsilon_{t+1}^{(10)}.$$

The regressions are estimated with monthly data from 1989m1-2011m4, so each month I forecast the excess return over the following 12 months. To deal with the overlapping nature of returns, *t*-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. I estimate these regressions with and without other forecasting variables identified in the literature on bond risk premia. Specifically, I control for the term spread following Campbell and Shiller (1991) and the first five forward rates following Cochrane and Piazzesi (2005).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
$c_t - y_{M,t}$	-2.571	-7.766	-6.867	2.444	-4.702	-5.158						
	[-1.26]	[-4.46]	[-3.29]	[0.72]	[-1.44]	[-1.51]						
$REFI_t$							-4.670	-6.772	-6.027	-2.264	-4.448	-4.432
							[-3.33]	[-6.31]	[-4.23]	[-1.12]	[-2.87]	[-2.59]
$DUR\_CNTRB_t^{MBS}$				15.470	7.686	4.624				7.682	7.378	5.551
				[2.72]	[1.90]	[1.17]				[2.04]	[2.83]	[2.37]
$y_t^{(10)} - y_t^{(1)}$		4.537			4.038			3.317			3.298	
		[4.70]			[3.92]			[4.55]			[4.65]	
$f_t^{(1)}$			-26.123			-23.254			-17.055			-14.383
			[-3.53]			[-3.91]			[2.27]			[-1.95]
$f_t^{(2)}$			131.748			113.968			94.061			74.597
			[3.09]			[3.50]			[2.43]			[2.03]
$f_t^{(3)}$			-318.861			-274.439			-246.166			-196.714
			[-3.34]			[-3.78]			[2.85]			[-2.42]
$f_t^{(4)}$			347.721			299.820			281.555			228.361
			[3.59]			[4.00]			[3.13]			[2.66]
$f_t^{(5)}$			-134.407			-116.017			-112.104			-91.553
			[-3.61]			[-3.89]			[3.15]			[-2.64]
Constant	4.303	-4.202	-1.734	-11.061	-10.900	-6.056	5.273	-0.210	0.632	-2.978	-8.102	-5.626
	[2.73]	[-1.87]	[-0.32]	[-1.94]	[-3.12]	[-1.13]	[4.17]	[0.13]	[0.13]	[0.73]	[2.27]	[-0.95]
Number of observations	268	268	268	268	268	268	256	256	256	256	256	256
R^2	0.05	0.42	0.50	0.19	0.45	0.51	0.17	0.42	0.51	0.21	0.46	0.52

Table 6

Forecasting excess bond returns, subsample results.

Regressions of 12-month excess returns on ten-year Treasuries on the effective duration of the MBS Index, the effective duration contribution of the MBS index, and aggregate refinancing incentives

$$rx_{t+1}^{(10)} = a + b \cdot DUR_t + \mathbf{c}'\mathbf{x}_t + \varepsilon_{t+1}^{(10)}.$$

For mortgage-backed securities (MBS) duration, DUR_t^{MBS}(effective) refers to option-adjusted effective duration which is available beginning in 1989. dur_t^{MBS}(Macaulay) refers to Macaulay duration-to-worst which is available beginning in 1976. A similar notation applies in the case of MBS duration contribution. Aggregate refinancing incentives $(c_t - y_{M,t})$, computed as the difference between the average MBS coupon in the Barclays Index and the Freddie Mac primary market 30-year mortgage rate, is also available beginning in 1976. The regressions are estimated with monthly data from 1976m1-2011m4, so each month I forecast the excess return over the following 12 months. To deal with the overlapping nature of returns, t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. I estimate these regressions with and without controlling for the term-spreads.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1989+	1989-1999	2000+	1989+	1976+	1976-1988	1989-1999	2000+
Panel A: Subsample results for MBS duration								
DUR_t^{MBS} (effective)	4.437	5.324	4.355					
	[6.48]	[3.33]	[9.44]					
dur_t^{MBS} (Macaulay)				5.968	1.505	-0.962	10.198	4.383
				[4.57]	[0.76]	[-0.27]	[3.05]	[5.14]
$y_t^{(10)} - y_t^{(1)}$	3.231	3.062	2.938	2.935	3.481	4.516	3.278	2.333
	[4.67]	[1.93]	[4.30]	[4.23]	[3.34]	[1.76]	[2.44]	[2.92]
Constant	-14.948	-19.279	-12.830	-27.100	-8.601	2.063	-48.689	-17.803
	[-4.58]	[-2.65]	[-4.73]	[-4.08]	[-0.93]	[0.11]	[-2.99]	[-4.01]
Number of observations	268	132	136	268	424	156	132	136
R^2	0.37	0.34	0.49	0.34	0.15	0.13	0.41	0.37
Panel B: Subsample results for MBS duration contribution								
$DUR\_CNTRB_t^{MBS}$ (effective)	14.629	19.993	11.886					
	[5.78]	[3.86]	[8.64]					
$dur\_cntrb_t^{MBS}$ (Macaulay)				14.551	6.733	9.220	33.513	11.891
				[3.80]	[2.18]	[1.71]	[2.96]	[5.76]
$y_t^{(10)} - y_t^{(1)}$	2.999	3.004	2.837	2.319	2.579	3.756	3.140	2.218
	[4.57]	[1.98]	[4.07]	[3.52]	[2.64]	[1.58]	[2.84]	[2.89]
Constant	-15.515	-21.018	-12.307	-20.371	-8.244	-9.578	-46.573	-16.979
	[-4.23]	[-3.04]	[-4.42]	[-3.39]	[-1.63]	[-1.84]	[-3.07]	[-4.22]
Number of observations	268	132	136	268	424	156	132	136
R^2	0.40	0.38	0.47	0.32	0.20	0.15	0.42	0.36
Panel C: Subsample results for mortgage refinancing incentive $(c_t - y_{M,t})$								
$c_t - y_{M,t}$				-7.766	-1.684	-3.455	-9.119	-6.717
				[-4.46]	[-1.50]	[-2.91]	[-4.24]	[-3.78]
$y_t^{(10)} - y_t^{(1)}$				4.537	4.531	8.205	6.066	3.828
				[4.70]	[3.94]	[3.87]	[3.54]	[4.14]
Constant				-4.202	-4.479	-15.428	-5.710	-3.209
				[-1.87]	[-1.79]	[-5.32]	[-1.70]	[-1.21]
Number of observations				268	424	156	132	136
R^2				0.42	0.19	0.30	0.48	0.33

Table 7

Using US mortgage-backed securities (MBS) duration to forecast foreign bond excess returns.

This tables presents univariate forecasting regressions of the form

$$rx_{c,t+1}^{FOR(n)} = a + b \cdot DUR_t + \varepsilon_{c,t+1}^{FOR(n)},$$

and multivariate regressions of the form

$$rx_{c,t+1}^{FOR(n)} = a + b \cdot DUR_t + c \cdot (y_{c,t}^{US(n)} - y_t^{US(1)}) + d \cdot (y_{c,t}^{FOR(n)} - y_t^{FOR(1)}) + \varepsilon_{c,t+1}^{FOR(n)}.$$

Using International Financial Statistics data on long-term government bond yields and short-term interest rates, I compute the excess returns on long-term government bond returns for the ten developed countries: Belgium, Canada, Denmark, France, Germany, Italy, Japan, Sweden, Switzerland, and the UK. Because the long-term yields are for coupon-bearing bonds, I follow Campbell (1999, 2003) and use the Shiller, Campbell, and Schoenholtz (1983) approximation to compute excess returns over short-term government bills (when available) or money market rates (when bills are not available). The regressions are estimated with monthly data from 1989m1-2011m4. To deal with the overlapping nature of the regressions, *t*-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. To assess the differential forecasting power between foreign bond and domestic bond excess returns, for each country I rerun the regressions with $rx_{c,t+1}^{FOR(n)} - rx_{c,t+1}^{US(n)}$ on the left-hand side. I then report the coefficient on DUR_t from this specification as well as the associated *t*-statistic, labeled as "Δ from USA." Panel A shows results for DUR_t^{MBS}, and Panel B shows results for $DUR\_CNTRB_t^{MBS}$.

	USA		Belgium		Canada		Denmark		France		Germany		Italy		Japan		Sweden		Switzerland		UK	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)
Panel A: Forecasting foreign bond excess returns using DUR_t^{MBS}																						
DUR_t^{MBS}	2.42	2.91	0.44	0.77	0.69	1.90	0.83	1.06	0.21	0.74	0.20	0.58	-0.44	-0.18	1.51	2.13	1.03	1.41	-0.14	1.37	0.92	1.50
	[3.53]	[5.81]	[0.52]	[1.44]	[0.86]	[4.36]	[1.19]	[1.78]	[0.31]	[1.59]	[0.23]	[0.88]	[-0.56]	[-0.20]	[1.21]	[2.49]	[1.07]	[1.73]	[-0.15]	[2.85]	[0.99]	[1.99]
$y_t^{US(n)} - y_t^{US(1)}$		1.94		1.37		0.56		1.63		1.52		1.93		1.04		2.10		1.38		2.12		1.49
		[3.13]		[1.80]		[0.88]		[2.28]		[2.00]		[2.75]		[0.86]		[2.74]		[1.56]		[3.19]		[2.12]
$y_t^{FOR(L)} - y_t^{FOR(S)}$				1.74		1.79		1.35		1.05		1.84		-1.03		2.15		1.39		1.80		0.79
				[2.21]		[2.58]		[2.29]		[1.53]		[2.51]		[0.79]		[0.98]		[1.95]		[3.31]		[1.18]
Constant	-4.64	-9.62	1.11	-4.82	1.19	-6.63	0.00	-4.59	1.81	-3.60	1.83	-4.82	4.05	2.52	-2.75	-10.59	-0.09	-5.18	2.73	-7.72	-0.53	-5.45
	[-1.74]	[-3.80]	[0.40]	[-2.16]	[0.48]	[-4.17]	[0.00]	[-1.82]	[0.80]	[-1.62]	[0.69]	[-2.00]	[1.58]	[0.71]	[-0.78]	[-3.95]	[-0.03]	[-1.84]	[0.98]	[-3.88]	[-0.19]	[-1.77]
Observations	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268
R^2	0.14	0.29	0.00	0.26	0.02	0.39	0.01	0.22	0.00	0.17	0.00	0.30	0.00	0.03	0.05	0.25	0.01	0.11	0.00	0.37	0.02	0.22
Δ from USA			-1.98	-2.10	-1.72	-1.07	-1.59	-1.79	-2.21	-2.22	-2.22	-2.30	-2.86	-3.12	-0.91	-0.79	-1.39	-1.51	-2.56	-1.74	-1.50	-1.40
			[-3.73]	[-4.88]	[-3.73]	[-2.66]	[-3.35]	[-3.40]	[-4.65]	[-4.98]	[-3.70]	[-4.28]	[-3.09]	[-3.46]	[-0.87]	[-0.88]	[-1.93]	[-2.13]	[-3.30]	[-2.78]	[-1.85]	[-1.89]
Panel B: Forecasting foreign bond excess returns using $DUR\_CNTRB_t^{MBS}$																						
$DUR\_CNTRB_t^{MBS}$	8.76	10.05	2.14	2.27	2.82	5.31	3.53	3.61	1.88	2.56	2.69	3.22	-2.14	-1.10	5.57	6.32	4.31	4.49	2.15	4.46	4.07	5.39
	[4.31]	[5.70]	[0.98]	[1.44]	[1.45]	[4.00]	[1.45]	[2.03]	[0.96]	[1.50]	[1.34]	[1.78]	[-0.94]	[-0.44]	[2.77]	[2.84]	[1.69]	[1.86]	[1.10]	[-3.05]	[1.76]	[2.37]
$y_t^{US(L)} - y_t^{US(S)}$		1.90		1.36		0.69		1.61		1.52		1.97		1.01		2.01		1.36		2.13		1.55
		[3.43]		[1.74]		[0.89]		[2.29]		[1.98]		[2.76]		[0.83]		[2.53]		[1.52]		[3.17]		[1.96]
$y_t^{FOR(L)} - y_t^{FOR(S)}$				1.70		1.57		1.31		0.96		1.78		-0.98		1.58		1.27		1.61		0.71
				[2.07]		[2.12]		[2.12]		[1.35]		[2.36]		[-0.75]		[0.69]		[1.66]		[3.01]		[1.06]
Constant	-6.01	-10.66	0.26	-4.58	0.47	-5.88	-1.05	-4.86	0.47	-3.81	-0.41	-6.34	4.89	3.12	-3.70	-9.53	-1.30	-5.15	-0.06	-7.77	-1.84	-6.30
	[-2.28]	[-4.05]	[0.10]	[-1.82]	[0.22]	[-3.27]	[-0.35]	[-1.83]	[0.19]	[-1.35]	[-0.18]	[-2.35]	[1.58]	[0.75]	[-1.79]	[-3.03]	[-0.41]	[-1.45]	[0.03]	[-3.84]	[-0.75]	[-1.83]
Observations	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268	268
R^2	0.19	0.33	0.01	0.26	0.03	0.37	0.02	0.22	0.01	0.17	0.02	0.32	0.01	0.03	0.07	0.23	0.02	0.11	0.01	0.38	0.04	0.23
Δ from USA			-6.63	-7.48	-5.94	-4.66	-5.23	-6.19	-6.88	-7.42	-6.07	-6.69	-10.91	-11.57	-3.20	-3.96	-4.45	-5.70	-6.61	-5.56	-4.70	-4.72
			[-5.29]	[-5.17]	[-5.57]	[-4.81]	[-4.06]	[-3.92]	[-5.38]	[-4.80]	[-4.70]	[-4.09]	[-3.88]	[-3.65]	[-1.33]	[-1.46]	[-3.04]	[-2.99]	[4.06]	[3.12]	[2.30]	[-2.26]

Table 8
Response of US long-term forwards to changes in short-term rates.
Daily regressions of changes in nominal and real forward rates on short-term nominal rates beginning in 1999, allowing for a differential response depending on mortgage-backed securities (MBS) negative convexity. MBS negative convexity is simply negative one times MBS convexity. Specifically, I estimate

$$\Delta f_t^{X(10)} = a + b \cdot \Delta y_t^{\$(2)} + c \cdot NCONV_t^{MBS} + d \cdot (\Delta y_t^{\$(2)} \times NCONV_t^{MBS}) + \Delta \varepsilon_t^{X(10)}$$

and

$$\Delta f_t^{X(10)} = a + b \cdot \Delta f_t^{\$(2)} + c \cdot NCONV_t^{MBS} + d \cdot (\Delta f_t^{\$(2)} \times NCONV_t^{MBS}) + \Delta \varepsilon_t^{X(10)},$$

for $X = \$$ and *TIPS*. *t*-statistics, based on Newey and West (1987) standard errors allowing for serial correlation at a lag of up to 20 business days, are shown in brackets.

	\$	*TIPS*	\$	*TIPS*	\$	*TIPS*	\$	*TIPS*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta y^{\$(2)}$	0.322	0.124	0.475	0.142				
	[2.67]	[1.86]	[3.99]	[1.98]				
$\Delta f^{\$(2)}$					0.264	0.074	0.364	0.084
					[3.33]	[1.41]	[4.97]	[1.66]
$NCONV_{MBS}$	0.000	-0.001			0.000	-0.001		
	[0.06]	[-0.58]			[0.13]	[-0.65]		
$NCONV^{MBS} \cdot (MV^{MBS}/MV^{AGG})$			0.002	0.000			0.001	-0.001
			[0.23]	[0.06]			[0.09]	[-0.16]
$\Delta y^{\$(2)} \times NCONV^{MBS}$	0.140	0.114						
	[1.74]	[2.51]						
$\Delta y^{\$(2)} \times NCONV^{MBS} \cdot (MV^{MBS}/MV^{AGG})$			0.107	0.278				
			[0.48]	[1.96]				
$\Delta f^{\$(2)} \times NCONV^{MBS}$					0.155	0.135		
					[2.94]	[4.00]		
$\Delta f^{\$(2)} \times NCONV^{MBS} \cdot (MV^{MBS}/MV^{AGG})$							0.247	0.354
							[1.77]	[3.74]
Constant	0.000	0.000	-0.001	-0.001	0.000	0.000	-0.001	0.000
	[0.01]	[0.19]	[-0.27]	[-0.28]	[0.02]	[0.19]	[-0.16]	[0.26]
R^2	0.21	0.12	0.21	0.12	0.31	0.19	0.31	0.19

Table 9
Implied variance of US long-term yields and mortgage-backed securities (MBS) negative convexity. Weekly regressions of swaption-implied yield variance on MBS negative convexity beginning in 1997

$$\tilde{\sigma}_t^2(\Delta y^{(10)}) = a + b \cdot \tilde{\sigma}_{t-1}^2(\Delta y^{(10)}) + c \cdot y_{t-1}^{(10)} + d \cdot NCONV_{t-1}^{MBS} + u_t.$$

MBS negative convexity is negative one times MBS convexity. I follow the empirical approach of Perli and Sack (2003). Specifically, the left-hand side variable, $\tilde{\sigma}_t^2(\Delta y^{(10)})$, is the residual from a regression of ten-year by three-month forward swaption-implied variance (the square of implied volatility) on ten-year by two-year forward implied yield variance. t-statistics, based on Newey and West (1987) standard errors allowing for serial correlation at a lag of up to 12 weeks, are shown in brackets.

	(1)	(2)	(3)	(4)
$\tilde{\sigma}_{t-1}^2(\Delta y^{(10)})$	0.941	0.941	0.939	0.939
	[49.81]	[50.40]	[49.61]	[49.86]
$y_{t-1}^{(10)}$		0.004		0.006
		[1.07]		[1.64]
$NCONV_{t-1}^{MBS}$	0.013	0.020		
	[2.14]	[2.29]		
$NCONV_{t-1}^{MBS} \cdot (MV_{t-1}^{MBS}/MV_{t-1}^{AGG})$			0.048	0.078
			[2.14]	[2.70]
Constant	-0.021	-0.051	-0.026	-0.072
	[-2.28]	[-1.62]	[-2.42]	[-2.31]
Number of observations	785	785	785	785
R^2	0.89	0.89	0.89	0.89

Table 10

The impact on mortgage-backed securities (MBS) duration on corporate bond and swap spreads

Panel A shows regressions of three-month changes in spreads over Treasuries on contemporaneous three-month changes in MBS duration:

$$\Delta SPREAD_t = a_1 + b_1 \cdot \Delta DUR\_CNTRB_t^{MBS} + \mathbf{c}_1' \Delta \mathbf{x}_t + v_t .$$

Panel B uses the current level of MBS duration to forecast changes in spreads over the next 12 months:

$$\Delta SPREAD_{t+1} = a_2 + b_2 \cdot DUR\_CNTRB_t^{MBS} + \mathbf{c}_2' \mathbf{x}_t + u_{t+1} .$$

Panel C uses the current level of MBS duration to forecast excess returns over duration-matched Treasuries over the following 12 months:

$$rx_{t+1}^{DUR-MATCH} = a_3 + b_3 \cdot DUR\_CNTRB_t^{MBS} + \mathbf{c}_3' \mathbf{x}_t + \varepsilon_{t+1} .$$

All regressions begin in 1999m1. To deal with the overlapping nature of the regressions, *t*-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to six lags in Panel A and up to 18 lags in Panels B and C. In Panel A, I control for changes in the term spread, the Chicago Board Options Exchange Market Volatility Index (VIX), and the past returns on the stock market, all of which Collin-Dufresne, Goldstein, and Martin (2001) show are useful for explaining changes in corporate credit spreads. In Panel B and C, I control for the initial level of the term spread, the VIX, and the initial level of the spread for the each asset. Corporate bond spreads in Panel A and B for investment-grade corporates, Aaa-rated corporates, and Baa-rated corporate are the (option-adjusted) spreads on Barclays indices over duration-matched Treasuries. The ten-year swap spread is from Bloomberg. The returns in Panel C are the excess returns on these indices over duration-matched Treasuries. For swaps, I do not have an exact measure of excess returns over duration-matched Treasuries. Thus, I use the Shiller, Campbell, and Schoenholtz (1983) approximation. I compute this excess return using the yield on ten-year swaps and the ten-year par Treasury yield from Gurknayak, Sack, and Wright (2007).

	IG corporate spread		Aaa corporate spread		Baa corporate spread		10-yr swap spread	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Contemporaneous 3-month changes in spreads over duration-matched Treasuries								
$\Delta DUR\_CNTRB_t^{MBS}$	-0.43	-0.16	-0.14	0.04	-0.72	-0.35	0.13	0.17
	[-3.56]	[-2.21]	[-1.92]	[0.50]	[-3.52]	[-2.75]	[3.40]	[5.22]
$\Delta(y_t^{(10)} - y_t^{(1)})$		0.05		0.04		0.06		-0.08
		[0.52]		[0.46]		[0.48]		[-2.98]
ΔVIX_t		0.03		0.03		0.03		0.00
		[2.78]		[2.40]		[2.70]		[1.98]
$MKTRF_t$		-0.01		0.00		-0.02		0.00
		[-2.40]		[1.14]		[-3.84]		[1.21]
Constant	0.01	0.03	0.00	0.00	0.01	0.04	-0.01	-0.01
	[0.19]	[0.79]	[0.06]	[0.17]	[0.13]	[0.94]	[-0.97]	[-1.10]
Number of observations	265	265	265	265	265	265	265	265
R^2	0.06	0.51	0.01	0.32	0.10	0.50	0.09	0.22
bps effect 1-σ DUR	-11.2	-4.0	-3.5	0.9	-18.8	-9.0	3.3	4.3

Table 10 (continued):

	IG corporate spread		Aaa corporate spread		Baa corporate spread		10-yr swap spread	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel B: Forecasting future 12-month changes in spreads over duration-matched Treasuries								
$DUR\_CNTRB_t^{MBS}$	0.85	0.63	0.31	0.30	1.18	0.92	-0.17	-0.16
	[2.68]	[3.72]	[1.46]	[2.81]	[2.93]	[4.07]	[-2.00]	[-2.87]
$y_t^{(10)} - y_t^{(1)}$		-0.12		-0.08		-0.11		-0.09
		[-1.44]		[-1.65]		[-1.23]		[-4.32]
VIX_t		0.00		0.01		0.01		0.00
		[0.15]		[0.81]		[0.76]		[0.44]
$SPREAD_t$		-0.47		-0.82		-0.61		-0.47
		[-1.89]		[-4.53]		[-2.88]		[-4.29]
Constant	-0.87	0.24	-0.33	0.34	-1.24	0.04	0.15	0.56
	[-2.52]	[0.68]	[-1.70]	[1.64]	[-2.61]	[0.11]	[1.55]	[4.25]
Number of observations	256	256	256	256	256	256	256	256
R^2	0.07	0.35	0.02	0.44	0.08	0.37	0.06	0.40
bps effect 1-σ DUR	-24.7	-18.4	-8.9	-8.6	-34.5	-26.8	4.9	4.7
Panel C: Forecasting 12-month excess returns over duration-matched Treasuries								
$DUR\_CNTRB_t^{MBS}$	-4.65	-3.59	-2.28	-1.58	-6.53	-5.10	1.20	0.49
	[-2.75]	[-3.87]	[-1.99]	[-2.31]	[-3.17]	[-4.66]	[2.08]	[1.02]
$y_t^{(10)} - y_t^{(1)}$		0.51		0.41		0.50		0.22
		[1.11]		[1.74]		[1.05]		[1.91]
VIX_t		0.03		0.05		-0.07		-0.01
		[0.25]		[0.87]		[-0.71]		[1.01]
$SPREAD_t$		2.49		0.15		3.44		3.86
		[1.74]		[0.16]		[3.09]		[4.48]
Constant	5.21	-0.57	2.38	-0.11	7.37	0.29	-1.10	-2.43
	[2.82]	[-0.30]	[2.41]	[-0.11]	[2.97]	[0.13]	[-2.26]	[-2.52]
Number of observations	256	256	256	256	256	256	256	256
R^2	0.07	0.32	0.06	0.13	0.08	0.38	0.05	0.32