Inflation Bets on the Long Bond

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The liquidity premium theory of interest rates predicts that the Treasury yield curve steepens with inflation uncertainty as investors demand larger risk premiums to hold long-term bonds. By using the dispersion of inflation forecasts to measure this uncertainty, we find the opposite. Since the prices of long-term bonds move more with inflation than short-term ones, investors also disagree and speculate more about long-maturity payoffs with greater uncertainty. Shorting frictions, measured by using Treasury lending fees, then lead long maturities to become overpriced and the yield curve to flatten. We estimate this inflation-betting effect using time variation in inflation disagreement and Treasury supply. (*JEL* G12)

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The liquidity premium theory of interest rates, the conventional explanation for why the Treasury yield curve slopes upward, predicts that when there is more uncertainty about inflation, the slope of the yield curve should, if anything, become steeper (Keynes 2006; Tobin 1958). Risk-averse investors with potential liquidity needs worry about having to sell during a bout of unexpected high inflation and depressed bond prices. They prefer, all else equal, short-term bonds, which are less sensitive to inflation and so they demand a risk premium to hold long-term bonds.[1] The higher yields at longer maturities generate higher expected returns to compensate investors to take on maturity

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[1] The premise of these models is that risk premium effects dominate convexity effects since convexity can also affect the slope of the yield curve.

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or duration risk. This risk premium accounts for the failure of the expectations hypothesis and gives rise to an upward-sloping yield curve.[2]

However, findings from recent studies suggest that this central prediction is unlikely to be true. First, Frazzini and Pedersen (2014) find that the Sharpe ratio of Treasury bonds monotonically declines with maturity, indicating that investors in long-term bonds are not compensated enough for taking duration risk. Second, and more broadly, Baker, Greenwood, and Wurgler (2003) argue that corporations take into account information about the flatness of the yield curve to time both the issuance and maturity of their debt. These studies, however, do not systematically attempt to understand the sources of potential mispricings and the role of inflation uncertainty in affecting the yield curve.

In this paper, we show that the slope of the term structure of expected bond returns does not increase with inflation uncertainty, and that, if anything, the opposite is true. Each month, we measure inflation uncertainty using the cross-sectional standard deviation of inflation forecasts made by U.S. households. Forecasts of inflation, and indeed of other macroeconomic variables, tend to become more dispersed during uncertain times (Cukierman and Wachtel 1979; Zarnowitz and Lambros 1987). The dispersion of these forecasts is thus a natural and real-time measure of inflation uncertainty. Our data, which cover the period from 1978 to 2012, come from the University of Michigan survey of consumer sentiment, which samples each month around 600 subjects from the general public. Among the various surveys available, Mankiw, Reis, and Wolfers (2004) point out that the Michigan series has the lowest forecast error and that surveys of households are more apt to pick up uncertainty compared to surveys of just professional economists.[3]

In Figure 1, we show that there is indeed dispersion in these Michigan inflation forecasts and that this dispersion varies substantially over time. Inflation disagreement is highest during the late 1970s. It also has mini-peaks in the early 1990s and most recently during the years following the Financial Crisis of 2008. To investigate the effect of inflation disagreement on bond excess returns, we compute, each month, the subsequent one-year holding period returns for Treasury bonds of various maturities, in excess of the one-year bond. We then split our time series into months of high inflation uncertainty (i.e., when uncertainty is in the top tercile of the in-sample distribution of inflation uncertainty) and months of low inflation uncertainty (i.e., when uncertainty is in the bottom tercile of the in-sample distribution of inflation uncertainty).

Figure 2 plots the average Treasury bond excess returns against the maturity of these bonds for low versus high inflation uncertainty or disagreement months.

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2 Other explanations, such as the preferred habitat hypothesis in which investors have preferred habitats in terms of maturities they want to own, yield a similar prediction about the shape of the yield curve (Culbertson 1957; Modigliani and Sutch 1966).

3 However, our results hold, as we verify below, when we use another survey, namely, the Livingston Survey of professional forecasters.

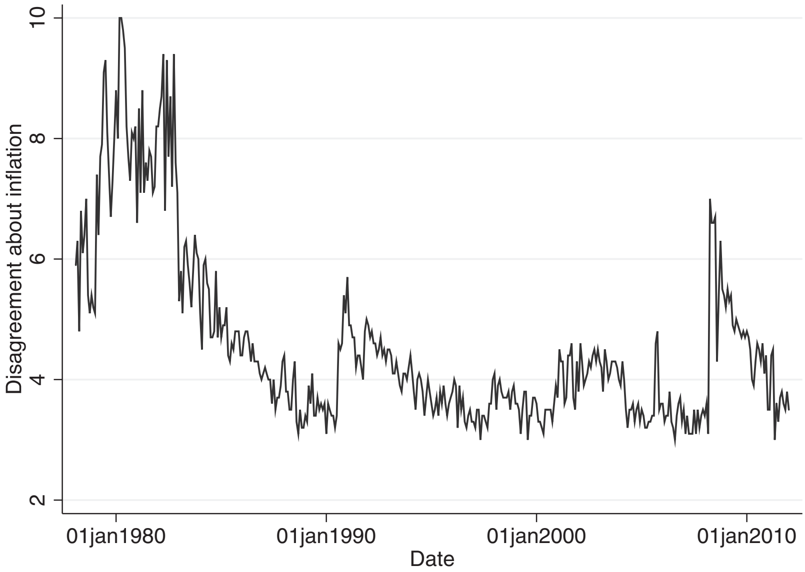

Figure 1
Time series of disagreement in inflation expectation
This figure plots the interquartile range in monthly inflation forecasts from Michigan Survey for 1978–2012. *Source:* Michigan Survey.

"x" represents the months in the bottom tercile of inflation uncertainty. "•" represents the months in the top tercile of inflation uncertainty. Notice that both curves are upward sloping, consistent with traditional theories of the yield curve based on liquidity premium or habitats. The higher expected returns obtained for longer maturity represents under these theories a risk premium to compensate investors for taking on duration risk. However, the slope of the curve represented by "•" is less steep than the slope of the curve represented by "x". That is, when inflation uncertainty is high and the risk premium ought to be the greatest, the term structure of Treasury returns is, if anything, less and not more steep.

To understand this relationship between inflation uncertainty and the slope of the term structure of Treasury returns, we propose another channel in the determination of interest rates. Times when uncertainty about inflation are high are also times when investors disagree about what inflation will be in the coming months. In other words, uncertainty among inflation forecasts can be taken as a proxy for actual heterogeneous expectations among bond investors in the same way that the literature has used disagreement among stock analyst forecasts as a proxy for disagreement about a stock's expected earnings (Diether, Malloy, and Scherbina 2002). Since a bond's sensitivity to inflation rises with maturity, a bond's sensitivity to disagreement about inflation also rises with maturity. Even small disagreements about the course of inflation are amplified into

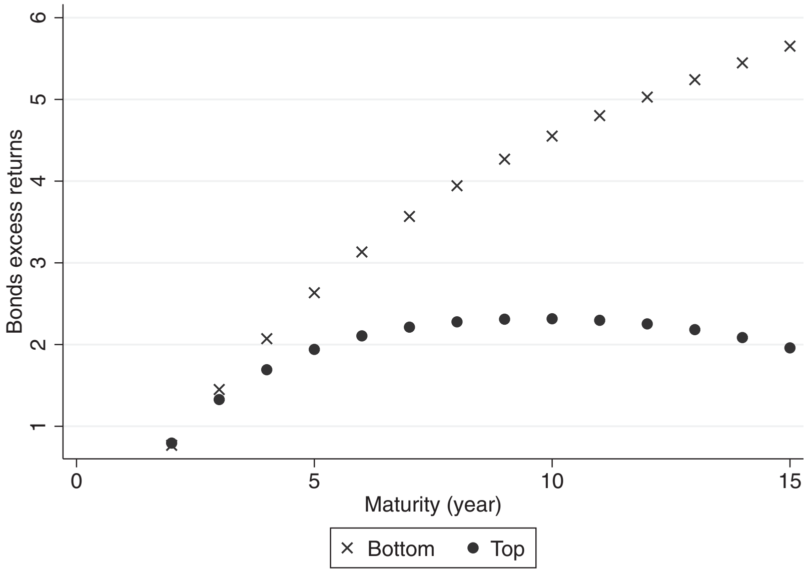

Figure 2
Bond excess returns for high versus low inflation disagreement months
This figure plots the average 1-year bond return for maturities of 2 to 15 years in excess of the 1-year Treasury bill for top and bottom tercile inflation disagreement months over the period of 1978–2012. Disagreement measure is the interquartile range of monthly inflation forecasts from the Michigan Survey.

large differences in expectations about the pay-offs of the long-term bonds. In contrast, even larger disagreements about inflation are dampened when it comes to the expectations of payoffs for short-term bonds.

So when uncertainty and disagreement rise, there is a new motive for trading in long-term bonds among investors.[4] Optimistic investors expecting low inflation now want to speculate and buy long-term bonds from pessimistic investors expecting high inflation who want to short. But some pessimists are likely to be short-sales constrained. While the Treasury market is often thought to be a venue where shorting frictions do not matter, we document in Section 3 that such frictions are consequential. Because of institutional restrictions, retail bond mutual funds, who own around 10% of the Treasury supply, do not short (Almazan et al. 2004; Koski and Pontiff 1999). We show that hedge funds, without such institutional restrictions, nonetheless, face nontrivial Treasury bond lending fees. As a result, long-term bonds will be held by the most optimistic investors in the market when inflation disagreement is large and short-sales constraints are binding. This then leads to more overpricing of long-term compared to short-term bonds and a flatter yield curve as a result.

4 Disagreement and speculative trade might arise for different reasons including differential interpretations (Harris and Raviv 1993; Kandel and Pearson 1995) and investor overconfidence (Odean 1999; Daniel, Hirshleifer, and Subrahmanyam 2001). See Hong and Stein (2007) for a review.

We term the amplification of inflation disagreement by bond maturity an inflation-betting effect.[5] Our work builds on that of Greenwood and Vayanos (2014), who develop a liquidity or habitat based theory of the term structure of interest rates. Like them, we use an overlapping generations model with mean-variance investors. Investors have access to a finite number of zero coupon bonds of different maturities, each with a positive supply, and a real asset with a deterministic real rate of return. The inflation rate, or more precisely the log growth rate of the price level, is modeled as an AR(1) process. Unlike their model, there are three groups of investors, optimists who expect the innovation in the inflation rate next period to be negative, pessimists who expect it to be positive and arbitrageurs. The optimists and pessimists can be thought of as bond mutual funds with retail investors who cannot short by charter. The arbitrageurs are hedge funds who can short for a fee. There is only one source of aggregate uncertainty: the realization of the innovation of the inflation rate. We solve for an equilibrium for bond prices, assuming a log-linearization of the investors' wealth process, and obtain the following key results.

Our model generates a key testable implication. When disagreement about inflation is low relative to the aggregate supply of bonds, short-sales constraints are nonbinding. Intuitively, a high aggregate supply of bonds will naturally depress bond prices due to the risk premium effect and lead even the most pessimistic of investors to own long-term bonds to share inflation risk. Risk premiums rise with maturity, yielding the standard prediction of an upward-sloping yield curve. More importantly, in this case, the slope of the yield curve increases with aggregate uncertainty, that is, inflation risk.

But when disagreement about inflation is high relative to the aggregate supply of bonds, short-sale constraints are binding. When short-sale constraints bind, the longer-maturity bonds are relatively more overpriced than short-term bonds. The slope of the yield curve is then flatter than when short-sale constraints are nonbinding. In other words, to understand why the relationship between inflation uncertainty and the slope of the term structure of bond returns in Figure 2 can be so flat, it is important to link the uncertainty of inflation forecasts to heterogeneous expectations and speculation on the part of bond investors and the aggregate supply of Treasuries.[6]

5 Our effect is related to an overpricing effect in stock markets. When there is high disagreement about aggregate market earnings, measured using the dispersion of security analysts' forecasts, Hong and Sraer (2016) show that the Security Market Line of the Capital Asset Pricing Model (Sharpe 1964) is too flat because beta amplifies disagreement about stock market earnings and short-sale constraints are more likely to bind for high beta stocks than for low beta ones. They term this a speculative beta effect.

6 The importance of supply for this binding short-sale constraint effect has been modeled in a static setting by Chen, Hong, and Stein (2002) and in a dynamic setting by Hong, Scheinkman, and Xiong (2006). High dispersion of security analysts forecasts also have been shown to forecast low returns in the cross-section of stocks (Diether, Malloy, and Scherbina 2002) and also low returns for the market in the time series (Yu 2011). Our main contribution is that we are first to test how this interaction of supply and disagreement forecasts asset returns.

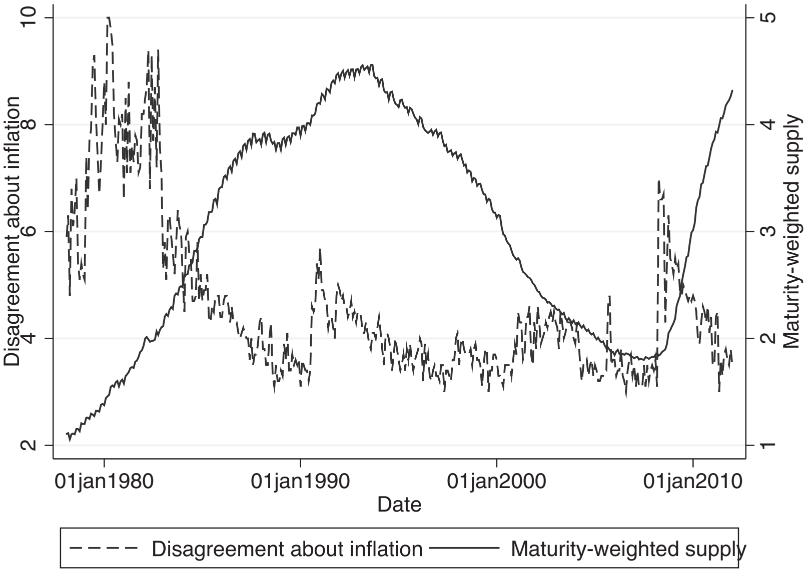

Figure 3
Time series of inflation disagreement and aggregate bond supply
This figure plots the inter-quartile range of monthly inflation forecasts from the Michigan Survey along with the ratio of maturity-weighted debt to GDP from Greenwood and Vayanos (2014) over the period of 1978–2012.

In our empirical work, we first establish that the supply of lendable Treasuries is not perfectly elastic by showing that bonds with higher lending fees underperform other bonds. We use the Markit lending fee data for Treasuries, which is available from 2011 to 2015. But given that shorting of Treasuries is likely to be even more difficult in the earlier periods, this finding confirms the premise of our model, namely that shorting frictions matter. As far as we know, this finding is new to the literature on the term structure of interest rates.

We then go on to study how Treasury excess returns vary with inflation uncertainty using panel data on Treasury bond returns and various disagreement measures. To identify our inflation-betting effect, we observe that the model predicts that short-sale constraints are more likely to be binding when disagreement is high *and* when supply of bonds is low. We follow Greenwood and Vayanos (2014) and report in Figure 3 the time series of the maturity-weighted supply of debt to GDP ratio. This is our empirical proxy for aggregate supply in our model. This ratio is rising from the late 1970s to the early 1990s and then declines through our remaining sample before peaking again with the unprecedented fiscal deficits since the Financial Crisis of 2008. Greenwood and Vayanos (2014) show that when aggregate supply is high, the yield curve is more upward sloping, consistent with risk aversion among bond investors.[7]

Since aggregate supply is initially rising until the 1990s with the fiscal expansion of President Reagan in the eighties and then declining with the surpluses in the President Clinton years, we can then use this aggregate supply variation to identify the inflation-betting effect on the shape of the yield curve. We thus run, for various maturities, a time-series regression of one-year holding period bond returns, in excess of the one-year Treasury Bond rate, on our inflation disagreement measure from Figure 1, the aggregate supply measure from panel A of Figure 3 and the interaction of these two variables. We use both linear specifications, as well as discrete specifications, where we split the time-series into terciles of inflation disagreement and aggregate bond supply, as shown in panel B of Figure 3.

Consistent with our model, disagreement about inflation expectation leads to lower expected excess returns for long-term bonds relative to short-term bonds when the aggregate supply of bonds is low relative to when it is high. This key result resists a variety of controls, including the other predictors of bond returns, such as the Cochrane and Piazzesi (2005) factor, business-cycle indicators, and subperiod breaks. We also conduct a series of robustness exercises, including different surveys to estimate disagreement, and different bond return series. We extend our analysis in a number of key dimensions.

First, we augment our regressions to include proxies for inflation risk so as to disentangle the effect of disagreement from inflation risk. Second, we verify the consistency of the term structure of inflation disagreement with the AR(1) process for inflation assumed in the model. Third, we relate our aggregate disagreement measure to trading volume and spreads in the Treasury market. Fourth, we consider interest rate disagreement. And, fifth, we connect our lending fee findings to inflation disagreement.

A number of recent papers point out that disagreement can affect the yield curve by affecting the volatility of bond prices in a dynamic setting even without frictions (Xiong and Yan 2010; Buraschi and Whelan 2012; Ehling et al. 2013). Notably, Ehling et al. (2013) focuses on how the level of bond yields is affected by investors' disagreement about inflation as opposed to our goal which is the slope of the term structure. There are no shorting frictions in their model in contrast with ours but they endogenize consumption in contrast with us. Their primary mechanism is that disagreement can lead to higher or lower real interest rates depending on income versus substitution effects because investors' consumption, and, hence, savings today are affected by their disagreement. This mechanism is initially presented in Gallmeyer and Hollifield

[7] Krishnamurthy and Vissing-Jorgensen (2012) find that a lack of Treasury supply raises the price of short-term debt as a safe asset. This effect works against us finding our disagreement and slope effect. Greenwood, Hanson, and Stein (2010) note the implications of Treasury supply for the issuance of corporate debt at various maturities. Malkhozov et al. (2016) find that mortgage supply can also affect the yield curve.

(2008), though their focus is on pricing the aggregate stock market. Ehling et al. (2013) find empirically that inflation disagreement leads to a higher level of bond yields across all bonds. Our primary mechanism is shorting frictions and supply affect the slope of the yield curve. Buraschi and Whelan (2012) take a similar framework as Ehling et al. (2013) and, again, in contrast with us has no shorting frictions, but consider a much broader set of state variables, including real GDP, over which investors might disagree. Their primary emphasis is that a host of disagreement variables about the real economy add incremental forecasting power relative to traditional term structure variables.

1. Model

We consider a discrete-time version of the bond pricing model in Greenwood and Vayanos (2014), where inflation is the sole source of risk for investors. There is an infinite number of periods $1,2,\dots,t,\dots,\infty$. Inflation follows an AR(1) process with long-term mean μ:

$$\tilde{\pi}_{t+1}=\mu+\rho(\pi_t-\mu)+\tilde{\epsilon}_{t+1}, \tag{1}$$

where $\mathbb{E}[\tilde{\epsilon}_{t+1}]=0$ and $Var[\tilde{\epsilon}_{t+1}]=\sigma_\epsilon^2$.[8]

The model features overlapping generations (OLG) of mean-variance investors who live for one period: generation t invests at t and consumes at $t+1$. Investors born at t can invest in a portfolio of zero-coupon bonds with K different maturities and in a real asset with a deterministic rate of return r in order to maximize their $t+1$ expected real wealth. In each generation, a fraction, α, of investors are Mutual Funds (MFs), and a fraction, $1-\alpha$, are Hedge Funds (HFs). MFs cannot short bonds. Shorting bonds is costly for HFs: to set up a short position x on a bond, these investors have to pay upfront a quadratic cost $\frac{c}{2}x^2$.[9] All investors can freely short the real asset.

Additionally, while HFs have homogeneous and rational beliefs about inflation, MFs are endowed with heterogeneous belief about the next period's expected value of inflation innovation: $\mathbb{E}^i[\tilde{\epsilon}_{t+1}]=\lambda^i$, $i\in(A,B)$.[10] MFs in group A (a fraction, 1/2, of the population of MFs) are optimists ($\lambda^A=-\lambda$) and investors in group B (a fraction, 1/2, of the population of MFs) are pessimists ($\lambda^B=\lambda$).[11] There is a deterministic supply of zero-coupon bonds of all maturity, which we call Q_t^k at date t for bonds of maturity k. Let Π_t be the price level at t.

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8 Throughout the model, variables with tildes are random variables, and we omit the tildes for their realizations.

9 The cost is quadratic. This is purely for analytical convenience. Our qualitative results hold as long as the cost is convex is the size of the short position.

10 For simplicity, we model the HFs as being arbitrageurs with well-calibrated beliefs. We can also introduce heterogeneous beliefs for HFs as well and obtain the same results.

11 Note that in our OLG setting, disagreement about the average ϵ is equivalent to disagreement about the long-run mean of inflation μ.

By definition: $\tilde{\Pi}_{t+1} = e^{\tilde{\pi}_{t+1}} \Pi_t$, that is, $\tilde{\pi}_{t+1}$ is the log-growth rate of the price index.

Generation-t investors are initially endowed at t with an exogenous real wealth W_t. Let V_{t+1} be their $t+1$ real wealth, which equals their $t+1$ consumption. $P_t^{(k)}$ is the price of a bond maturing in k periods at date t and $x_{t,i}^{(k)}$ is the number of bonds of maturity k held by investors in group i at date t. The $t+1$ real wealth of HFs (indexed by a for arbitrageur) is given by

$$\tilde{V}^a_{t+1} = \frac{\sum_{k=2}^K x_{t,a}^{(k)} \tilde{P}_{t+1}^{(k-1)} + x_{t,a}^{(1)}}{\tilde{\Pi}_{t+1}} + \left(W_t - \frac{\sum_{k=1}^K x_{t,a}^{(k)} \left(P_t^{(k)} + \frac{c}{2} x_{t,a}^{(k)} \mathbb{1}_{x_{t,a}^{(k)}<0} \right)}{\Pi_t} \right) (1+r). \tag{2}$$

The $t+1$ real wealth of MF in group $i \in \{A, B\}$ is given by:

$$\tilde{V}^i_{t+1} = \frac{\sum_{k=2}^K x_{t,i}^{(k)} \tilde{P}_{t+1}^{(k-1)} + x_{t,i}^{(1)}}{\tilde{\Pi}_{t+1}} + \left(W_t - \frac{\sum_{k=1}^K x_{t,i}^{(k)} \left(P_t^{(k)} \right)}{\Pi_t} \right) (1+r) \quad \text{and} \quad x_{t,i}^{(k)} \geq 0 \tag{3}$$

In what follows, we normalize r to 0 without loss of generality. We define the yield on a bond of maturity k at date t as: $k y_t^{(k)} = -\log\left(P_t^{(k)}\right)$. The optimal investment strategy of generation-t investors in group i is given by the following objective, where γ is investors' risk tolerance:

$$\max_{\left(x_{t,i}^{(k)}\right)} \mathbb{E}_t^i[\tilde{V}^i_{t+1}] - \frac{1}{2\gamma} \mathrm{Var}_t^i[\tilde{V}^i_{t+1}], \tag{4}$$

where γ is the aggregate risk aversion of each group of investors.

The following theorem characterizes equilibrium holding and prices as a function of disagreement λ:

Theorem 1 (Disagreement and Expected Bond Returns). Define $\Omega_t = \left(\sum_{k=1}^K \frac{1-\rho^k}{1-\rho} Q_t^{(k)} \right)$ and $\theta = \frac{\frac{\alpha}{2}}{1-\frac{\alpha}{2}} \in (0,1)$. Three cases arise:

1. When $\lambda < \frac{\sigma_\epsilon^2}{\gamma} \Omega_t$, all investors hold a long portfolio of bonds. The yield and expected 1-period holding return of a bond of maturity (k) are given, respectively, by

$$\begin{cases} y_t^{(k)} = \mu + \left(\dfrac{1-\rho^k}{k(1-\rho)} \right) \rho(\pi_t - \mu) + \dfrac{1}{k} \left(\displaystyle\sum_{l=1}^k \dfrac{1-\rho^l}{1-\rho} \right) \dfrac{\sigma_\epsilon^2}{\gamma} \Omega_t \\ \mathbb{E}[\tilde{R}_t^{(k)}] = \mu + \rho(\pi_t - \mu) + \left(\dfrac{1-\rho^k}{1-\rho} \right) \dfrac{\sigma_\epsilon^2}{\gamma} \Omega_t. \end{cases} \tag{5}$$

2. When $\frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t > \lambda \geq \frac{\sigma_\epsilon^2}{\gamma}\Omega_t$, pessimist MFs are sidelined from the bond market but optimist MFs and HFs still hold a long portfolio of bonds of all maturities. In this case, the yield and expected 1-period holding return of a bond of maturity (k) are given, respectively, by

$$\begin{cases} y_t^{(k)} = \mu + \left(\dfrac{1-\rho^k}{k(1-\rho)}\right)\rho(\pi_t-\mu) + \dfrac{1}{k}\left(\displaystyle\sum_{l=1}^{k}\dfrac{1-\rho^l}{1-\rho}\right) \\ \qquad \left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + \theta\left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t - \lambda\right)\right) \\ \mathbb{E}[\tilde{R}_t^{(k)}] = \mu + \rho(\pi_t-\mu) + \left(\dfrac{1-\rho^k}{1-\rho}\right)\left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + \theta\left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t - \lambda\right)\right). \end{cases} \tag{6}$$

3. When $\lambda > \frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t$, pessimist MFs are sidelined from the bond market and HFs hold a short portfolio of bonds of all maturities. Optimist MFs hold a long portfolio of all bonds. Define $\Theta = \sum_{k=1}^{K}\left(\frac{1-\rho^k}{1-\rho}\right)^2$. The yield and expected 1-period holding return of a bond of maturity (k) are given respectively by:

$$\begin{cases} y_t^{(k)} = \mu + \left(\dfrac{1-\rho^k}{k(1-\rho)}\right)\rho(\pi_t-\mu) + \dfrac{1}{k}\left(\displaystyle\sum_{l=1}^{k}\dfrac{1-\rho^l}{1-\rho}\right) \\ \qquad \left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + \theta\left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t - \lambda\right)\right)\left(\dfrac{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}\right) \\ \mathbb{E}[\tilde{R}_t^{(k)}] = \mu + \rho(\pi_t-\mu) + \left(\dfrac{1-\rho^k}{1-\rho}\right)\left(\dfrac{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}\right) \\ \qquad \left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + \theta\left(\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t - \lambda\right)\right). \end{cases} \tag{7}$$

Proof. See Appendix A.1. ■

The structure of the equilibrium crucially depends on the relative magnitudes of disagreement about inflation (λ) and the aggregate supply of bonds (Ω_t).[12]

[12] If ρ is close to one, then $\Omega_t = \left(\sum_{k=1}^{K}\frac{1-\rho^k}{1-\rho}Q_t^{(k)}\right) \approx \left(\sum_{k=1}^{K} kQ_t^{(k)}\right)$. Notice that the relationship also depends on the risk tolerance of investors γ and the variance of inflation σ_ϵ. We focus on discussion on λ and Ω_t as these two time-varying variables are the key to our empirics. We view γ as fixed through time. We will discuss how to potentially disentangle σ_ϵ^2 below.

When disagreement about inflation is low relative to the aggregate supply of bonds, short-sale constraints are nonbinding. Intuitively, a high aggregate supply of bonds will naturally depress bond prices due to the risk premium effect and lead even the most pessimistic of investors to own long-term bonds to share inflation risk.

When disagreement is at moderate levels, relative to a fixed aggregate supply (i.e., Case 2), the pessimistic MFs hit binding short-sale constraints first and are sidelined. But the HFs, whose beliefs are between those of the pessimistic and optimistic MFs, are still long bonds. But when disagreement is at high levels (i.e., Case 3), the HFs start shorting for a fee c, while the optimistic MFs are of course the only investors now long the bonds.

Several natural comparative statics emerge from Theorem 1, which we collect in the following corollary to better understand the pricing implications of our equilibrium.

Corollary 1. Expected 1-period holding returns of bonds are (1) weakly increasing with the weighted-supply measure Ω_t, (2) weakly decreasing with inflation disagreement λ, and (3) weakly decreasing with the cost of short selling c. Additionally, the negative effect of disagreement on bond returns is (1) stronger for bonds of longer maturity and (2) stronger when the weighted-supply measure Ω_t is low.

Proof. See Appendix A.2. ■

First, consider variations in weighted-bond supply Ω_t for a given λ. When supply is high enough that all investors are long all bonds, an increase in Ω_t increase the quantity of inflation risk that investors have to bear, so that the returns on bonds of all maturities increase. This is the standard liquidity premium effect. When supply declines, so that pessimistic MFs are now sidelined from the bond market but HFs are still long bonds, bonds become overvalued. In this regime, an increase in weighted supply leads to an increase in the risk premium required by optimistic MFs and HFs to hold these bonds in equilibrium. Additionally, an increase in weighted supply reduces the mispricing on all bonds as it decreases the speculative demand for bonds by optimistic MFs. Finally, when supply becomes so small that HFs are short bonds, an increase in weighted supply Ω_t has a similar effect on bond expected returns – it increases the risk premiums required in equilibrium and decreases mispricing by reducing the speculative demand by optimistic MFs. Since returns are continuous in Ω_t, it directly follows that returns are strictly increasing with Ω_t, the bond weighted supply.

An increase in inflation disagreement λ for a given weighted-supply Ω_t leads to a decrease in returns. Of course, when disagreement is so low that all investors are long bonds, returns do not depend on λ. When λ becomes large enough that pessimistic MFs are sidelined from the bond market, but HFs are still long bonds, bonds become mispriced. The extent of this mispricing depends on the strength of the speculative demand by optimistic MFs, which is directly increasing with disagreement λ. The same reasoning applies when optimistic MFs remain the only investors long bonds in the market. As with aggregate supply, the continuity of returns with disagreement λ ensures that expected returns are weakly increasing with λ.

The effect of the short-selling cost c on returns is also intuitive. As long as HFs are long bonds, short-selling costs have no effect on bond returns. However, when disagreement about inflation is high enough that HFs end up shorting bonds, an increase in c decrease the arbitrage activity of HFs, which lead to an increase in bond prices and thus to a decline in bond expected returns.

Since long-maturity bonds are more sensitive to inflation, there is naturally more disagreement about the pay-offs of long-maturity bonds compared to short-maturity bonds. As a result, there is more speculative demand for the k-maturity bond than the $(k-1)$-maturity bond. Combined with the short-selling restriction, this means that the (k)-maturity bond is more overpriced and has lower expected returns than the $(k-1)$-maturity bond.

Corollary 1 also shows that the negative effect of disagreement on bond expected returns is stronger when weighted supply is low. This comparative static is a simple consequence of the fact that a decrease in weighted supply makes it more likely that the short-sale constraint binds for pessimist MFs, thus making mispricing of bonds more likely.

We now present the main empirical predictions that we derive from our theoretical analysis and that serve as a basis for our empirical investigation in the following section. These predictions essentially are derived from Corollary 1.[13] First, we note note that our entire analysis crucially relies on the importance of short-selling costs inhibiting arbitrageurs in the Treasury market. To investigate the actual role of shorting frictions in the Treasury market, Prediction 1 considers the effect of shorting costs on bond excess returns:[14]

Prediction 1. Bonds with higher shorting costs are more overpriced and have lower expected excess returns.

Prediction 2 is straightforward from our discussion of Corollary 1, namely inflation disagreement leads to more overpricing and lowers bond returns when aggregate supply of Treasuries is low:

Prediction 2. Expected bond excess returns decrease with disagreement when the maturity-weighted supply is low.

[13] In these predictions, we consider the case ρ is close to one, so that the weighted supply Ω_t is approximately equal to the maturity-weighted supply of Greenwood and Vayanos (2014).

[14] We have data on shorting costs at the bond level for a limited period of time (2010-2012). As a result, we cannot test predictions that would link bond excess returns, shorting fees, and time-series variation in inflation disagreement. This is why Prediction 1 is limited to the effect of shorting costs on unconditional bond returns.

Since the effect of inflation disagreement is larger for long-maturity bonds than short-maturity bonds, Prediction 3 relates the slope of the term structure of expected bond returns with inflation disagreement when the aggregate supply of Treasuries is low:

Prediction 3. The term structure of expected bond returns is flatter when inflation disagreement is high and when maturity-weighted supply is low.

Note that in all our empirical tests, we will calculate excess returns relative to a variety of well-known Treasury bond factors in the literature, as well as adjust these returns for liquidity, as measured by bid-ask spreads. Note also that in the Online Appendix, we extend our OLG model to allow for time-varying disagreement λ. The time-varying disagreement model introduces an additional effect – long bonds are exposed to disagreement risk and hence could receive a risk premium. We show in the Online Appendix that, as long as disagreement is persistent enough, the empirical predictions are the same as the constant-disagreement model.

2. Data and Variables

We use survey data of inflation forecasts from the Michigan Survey, which are available monthly from 1978 to 2012. Each month, we calculate $Disagreement_{t-1}$ as the inter-quartile range of 1-year inflation forecasts.[15] In our robustness checks, we also use the Livingston Survey from 1952 to 2012. Unfortunately, this survey only samples semiannually, in the months of June and December. As we show below, we have far less informative variation in our right-hand side variable $Disagreement_{t-1}$ when using the Livingston Survey than the Michigan Survey. We want our baseline series to capture as much variation in disagreement as possible, both across forecasters at a point in time and across time. This is why we make the Michigan Survey our baseline sample.

Following Greenwood and Vayanos (2014), the monthly series of supply of Treasuries is the maturity-weighted-debt-to-GDP ratio

$$Supply_t = \frac{\sum_{0<\tau\leq 30} D_t^{(\tau)}\tau}{GDP_t},$$

computed by multiplying the payments $D_t^{(\tau)}$ for each maturity τ times τ, summing across maturities, and scaling by GDP. $D_t^{(\tau)}$ includes both coupon and principal payments. The data come from CRSP Treasury database. We exclude tips, flower bonds, and other bonds with special tax status. Our results

[15] We follow Mankiw, Reis, and Wolfers (2004) in focusing on the interquartile range as a more robust statistic for disagreement than the standard deviation. We have, however, also checked our results using standard deviation and find largely similar ones.

are largely the same when we also include repos into the Treasury Supply (see Online Appendix Table A6). The repo-augmented supply equals supply multiply 1 + Total Size of Repo/Total Market Capitalization of Treasuries. The data on the size of repos are from the New York Federal Reserve Web site and start from 1998. Before 1998, we set the size of repos to 0.

Our bond return data are from Gurkaynak, Sack, and Wright (2007) and are available on the Federal Reserve Bank website, which provides the returns of individual bonds at various maturities. We equal-weight these returns to then analyze each month the 1-year holding period returns for the 2-, 3-, 4-, 5-, 5/10-, and 10/15-year bonds in excess of the 1-year bond.[16] We define $\tilde{R}_t^{(k)}$ as the one-year holding period return of the k-th maturity bond in excess of the 1-year maturity bond.

The summary statistics for these variables are reported in Table 1. $Disagreement_{t-1}$ has a mean of 4.6% and a standard deviation of 1.5%. The time-series of $Disagreement_{t-1}$ is shown in Figure 1. For example, in December 2012, the 25th percentile forecast of 1-year inflation is 1.5%, while the 75th percentile is 5.2%, so that $Disagreement_{t-1}$ for December 2012 is 3.7%, which can be seen as the last observation in Figure 1. Notice that disagreement has varied significantly over our sample period. It starts at around 6% in 1978, but this monthly series fluctuates quite a bit, dropping to less than 5% in the middle of 1978 and reaching a high of 10% in 1981. There is a precipitous drop in inflation disagreement in the mid-eighties followed by a much more gradual march downwards in inflation disagreement until the early nineties. Then disagreement jumps again in the early nineties to levels that were as high as parts of the late seventies. The decade between the mid-nineties and the mid-2000s was a tranquil period with disagreement as low as 3%. But this changes during the financial crisis after the Lehman Brothers bankruptcy in 2008, and disagreement shoots up to near 7%. More recently, disagreement has fallen to the levels of the tranquil period of the mid-nineties to mid-2000s.

We also report the summary statistics for the Livingston Survey in Table 1. Notice that the mean of the disagreement variable is far lower for Livingston than Michigan. It is 1.03 as opposed to 4.6 for the Michigan Survey. Notice also that the standard deviation of the Livingston disagreement series is far lower, at 0.45 compared to 1.5. This is a reflection of the monthly sampling of Michigan that allows us to capture disagreement not possible in the Livingston series. Indeed, we plot in Figure A1 in the Online Appendix the Livingston disagreement series and we see much lower disagreement and much less variation in this series compared to the Michigan series over the years that overlap. A more careful inspection of the figure also reveals that Michigan

[16] We obtain similar results using the CRSP Fixed Term Index and Fama bond price series. Since we are trying to analyze the term structure, we get much more long-end maturities from the Fed series than the Fama Bonds or the CRSP Fixed Term Indices. The disadvantage is that there is interpolation on the yields of some of the long-end maturities in the Fed series. As a result, we also consider a number of robustness checks using the other two series.

Table 1
Summary statistics

Panel A. Time-series variables

	Mean	SD	p(10)	p(25)	p(50)	p(75)	p(90)	Obs.
$Disagreement_{t-1}$ (Michigan)	4.60	1.50	3.30	3.50	4.10	4.90	7.10	407
$Disagreement_{t-1}$ (Livingston)	1.03	0.45	0.55	0.69	0.92	1.26	1.59	407
$Supply_{t-1}$	2.99	1.02	1.72	2.04	3.06	3.91	4.29	407
CP Factor	0.45	2.75	−2.89	−1.34	0.46	1.88	3.64	408
LN	1.37	1.86	−0.77	0.04	1.15	2.62	3.85	311
$\widehat{Slope}$	0.28	1.19	−1.13	−0.38	0.27	0.97	1.74	408
Excess returns:								
2-year	0.73	1.86	−1.44	−0.36	0.71	2.09	2.94	408
3-year	1.32	3.43	−2.78	−0.75	1.34	3.58	5.56	408
4-year	1.84	4.80	−3.85	−1.19	1.93	5.18	7.85	408
5-year	2.29	6.05	−5.08	−1.51	2.68	6.31	9.66	408
5/10-year	3.30	9.42	−8.93	−2.67	3.94	8.80	14.57	408
10/15-year	4.29	14.78	−13.91	−4.13	5.36	12.85	22.62	408

Panel B. Cross-sectional shorting variables

	Mean	SD	p(10)	p(25)	p(50)	p(75)	p(90)	Obs.
Fee	0.034	0.072	−0.036	−0.011	0.024	0.060	0.106	36
IssueSize	32.8	16.7	15.6	21.8	30.5	38.7	48.4	36
ValueShorted	1.24	1.27	0.26	0.50	0.87	1.53	2.52	36
Inventory	2.95	2.60	0.81	1.34	2.28	3.68	5.44	36

Panel A of this table presents summary statistics for the time-series variables used in our analysis. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available from the Federal Reserve Bank Web site. Disagreement (Michigan) is the monthly interquartile range of consumers forecast for the next year's inflation rate in the Michigan Survey. Disagreement (Livingston) is the monthly interquartile range of individual forecast for the next year's inflation rate in the Livingston Survey. Supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). CP Factor is the factor from Cochrane and Piazzesi (2005). LN is the bond return factor $F6$ in Ludvigson and Ng (2009). Excess Returns are the 1-year holding period return of bonds in excess of the one year bond returns for 2, 3, 4, 5, 5- to 10- and 10+ years of maturity. Slope is the slope coefficient in the monthly cross-sectional regression of bond excess return on bond maturity. Panel B of this table presents the time-series average of the monthly cross sectional mean and standard deviation of variables related to shorting individual bonds. These variables include shorting fee (Fee, in %, winsorized at the 1st and 99th percentiles), total size of a bond issue (IssueSize, in billion \$), total value of the bond issue already shorted (ValueShorted, in billion \$), and value of the bond issue available to lend (Inventory, in billion \$). The shorting data are from Markit for the sample period January 2010 to December 2012.

indeed picks up more variations in disagreement, including during the early nineties recession, which triggered disagreement on inflation expectations. These variations are not in the Livingston series.

The $Supply_{t-1}$ variable has a mean of 2.99 and a standard deviation of 1.02. The time series of $Supply_{t-1}$ is shown in Figure 3, along with the $Disagreement_{t-1}$ series. Aggregate bond supply starts at a low ratio of around 2 and gradually rises until the early nineties before beginning to fall until at the onset of the financial crisis in 2008; at this point the aggregate bond supply picks up again.

We construct the monthly CP_{t-1} factor following Cochrane and Piazzesi (2005).[17] The CP factor has a mean of 0.45 and a standard deviation of 2.75. The time series of CP is plotted along with $Disagreement_{t-1}$ in Figure 4. We

[17] To compute the CP_{t-1} factor, we first regress the average excess return on 2-, 3-, 4- and 5-year bonds on the 1-year yield and the 2-, 3-, 4- and 5-year forward rate using Fama-Bliss discount bonds. We run this regression over the same sample period than Cochrane and Piazzesi (2005) and use the predicted value over the entire

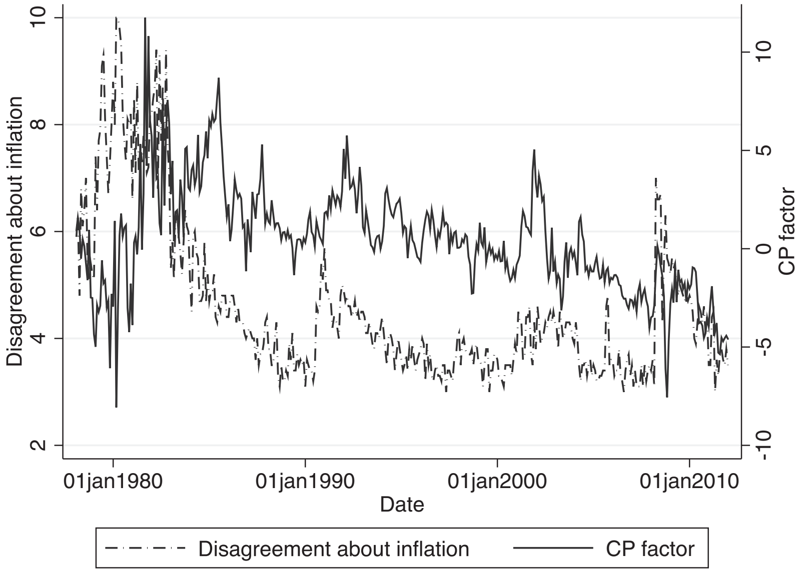

Figure 4
Time series of inflation disagreement and Cochrane-Piasezzi factor
This figure plots the interquartile range of monthly inflation forecasts from the Michigan Survey along with the Cochrane and Piazzesi (2005) factor over the period of 1978–2012.

can see on this figure that these two series have a somewhat positive correlation. We will hence think of the *CP* factor as a control variable to soak up omitted variables related to risk premiums in bond markets in our regressions.

In addition to the *CP* factor, our analysis also controls for business-cycle variables, such as the Ludvigson and Ng (2009) macro factor (*LN*) and the NBER recession dates.[18] We also control for the aggregate trading volume in the Treasury market (Volume), which we construct from GovPX data between 1991 and 2001 and from SIFMA after 2001. SPREAD is the bid-ask spread divided by mid-quote and is from CRSP. It is constructed by first averaging across bonds and then averaging across days in the previous month.

The summary statistics for the excess returns of bonds of various maturity are also reported in Table 1. The mean 1-year holding period return in excess of the 1-year bond rises from 0.73% for the 2-year maturity to 4.29% for the 10/15 year maturity. The standard deviations also rise from 1.86% to 14.78%. The yield curve is, on average, upward sloping, comparable to the results found in Fama (1984).

sample period as the CP factor. The results are similar if we instead run the initial regression over our entire sample period.

18 Ludvigson and Ng (2009) form factors from a large dataset of 132 macroeconomic indicators to conduct a model-free empirical investigation of reduced-form forecasting relations suitable for assessing more generally whether bond premiums are forecastable by macroeconomic fundamentals.

In addition to excess bond returns being our dependent variables of interest, we will also use the $Slope_t$ of the term structure of bond returns each month. We run a cross-maturity regression each month t to obtain an estimate of $Slope_t$ (i.e., $\hat{Slope}_t$):

$$\tilde{R}_t^{(k)} = \delta_t + Slope_t \times k + \epsilon_{k,t}, \tag{8}$$

where k is the maturity of the bond at t.[19]

Finally, in addition to inflation disagreement, supply and traditional bond market variables, we also obtain from Markit a dataset on lending fees for Treasuries over the 2010–2012 period. The Markit database covers the vast majority of lending transactions in the Treasuries market. The database has a structure similar to their well-known equities lending database (see, e.g., Davolio 2002), but has not been used previously in the literature. Summary statistics on lending fees are reported in panel B of 1, which we discuss below.

3. Shorting Frictions in the Treasuries Market

As we mentioned at the outset, there are two sources of short-sale constraints in the Treasury bond market. First, a large fraction of retail bond mutual funds are prohibited from shorting by charter.[20] As a result, mutual bond funds who are pessimistic about inflation and could short say the 30-year Treasury bond mostly sit on the sidelines. Pessimistic hedge funds have the ability to short and thus fill in for the pessimistic mutual funds. But as we document more extensively in this section using the Markit database, and consistent with earlier results in Duffie (1996), these hedge funds would face significant shorting costs.

Panel B of Table 1 reports summary statistics of the Markit Treasury securities lending data and document the size of shorting fees for Treasuries. Each month, we calculate the mean and standard deviation of the lending fee and other variables of interest from the lending market. We then report the time series average of these cross-sectional means and standard deviations. The summary statistics are similar to those reported in equity lending markets and corporate bond lending markets. The average lending fee for transactions in the data is around 4 basis points with a standard deviation of 7 basis points. The average short interest in a bond is 1.14 billion dollars. The average bond *IssueSize* is 33 billion dollars. Hence, the short ratio, defined as short interest (*ValueShorted*) over *IssueSize* is low at around 3.52%. This result is analogous to what is typically found on equity markets. The *Inventory* of shares available to be lent is 2.95 billion dollars and the utilization rate is 41.37% on average.

[19] Note that the estimate $\hat{Slope}_t$ is the excess return on a "carry" strategy, long (short) bonds that have longer (shorter) maturity than the average bond in the sample and where the portfolio weights are proportional to the relative maturity of each bond in the portfolios.

[20] Consistent with the lack of short interest in fixed income by retail and even institutional investors, among the Top 100 ETFs based on asset size, there are 17 fixed income ETFs and only one of these is a short fund. This fund only represents 3% of the total fixed income ETF assets.

Importantly, note that these measures of shorting frictions in the Treasury market are most likely a lower bound on the actual frictions prevailing on this market. Our data on shorting fees only cover a recent period (2010–2012), where shorting frictions were undoubtedly much more pronounced than in the earlier parts of our sample. Additionally, there are well-known limits to arbitrage that restricts the extent of shorting that arbitrageurs can do and that we do not document explicitly.

To further assess the relevance of shorting frictions in the Treasury market, we test the prediction that high lending fees for a long-maturity bond in month $t-1$ predicts underperformance for long-maturity bonds in month t (Prediction 1). This prediction is analogous to the standard test in equity markets, where high lending fees at the stock level predicts low stock returns (Jones and Lamont 2002). In other words, this analysis will establish that the supply of lendable Treasuries is not perfectly elastic. To implement this test, we divide our sample of bonds into short-maturity bonds of up to 5 years in maturity and long-maturity bonds defined as greater than 5 years. We then simply run a Fama-MacBeth regression on these two subsamples, where the monthly regression projects bonds excess returns on the aforementioned characteristics from the lending market as well as a dummy variable *OnRun*, which takes the value of 1 if the bond is on-the-run and zero otherwise.[21] The results are presented on Table 2, where columns (1) to (6) show the results for short-maturity bonds and column (7) to (12) show the results for long-maturity bonds. The estimated effect of short fees on future bond excess returns is negative in both sample, although the coefficient estimated on the sample of long-maturity bonds is markedly larger and statistically significant, which is consistent with our prediction. The point estimate in Column (12) implies that a one standard deviation increase in Fee (around 7 bps) is associated with a lower bond excess return next month of -21 bps. Since the average 1-month return of long maturity bonds is about 0.86 bps, this effect is economically significant.

Table 3 re-estimates the relationship between lending fee and bond excess returns by combining both short and long maturity bonds. To control for bonds maturity, we simply add the log of time to maturity as an extra control (*LogMat*). Columns (1) to (6) report the results of this estimation by showing how the estimated effect of lending fees on bond returns is affected as we add in more controls. Column (6), which contains all the additional control variables, reports a coefficient estimate for $Fee_{i,t-1}$ of -1 with a t-statistic of -1.7. This estimated effect implies that a one standard deviation increase in *Fee* is associated with a decrease in average bond returns of about -7 bps. Again, given an average monthly bond return of about 15 bps in sample, this point estimate represents an economically large decrease in bond returns for high fee bonds. The coefficient in front *LogMat* is as expected positive, implying an upward sloping term

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[21] On-the-run bond is defined as the most recently issued bond for a given maturity. Issue date is the TDATDT variable from CRSP.

Table 2
Bond return and shorting fee, by maturity bin

$r_{i,t} = b_1 Fee_{i,t-1} + b_2 OnRun_{i,t-1} + b_3 IssueSize_{i,t-1} + b_4 ValueShorted_{i,t-1} + b_5 Inventory_{i,t-1} + b_0 + \epsilon_{i,t}$

	(0,5y]						(5,30y]					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
b_1	−0.097	−0.13	−0.11	−0.16	−0.075	−0.13	−3.4*	−3.7*	−3.5*	−3.4**	−3.5*	−3.3**
	(−0.97)	(−1.1)	(−1.2)	(−1.4)	(−0.78)	(−1.4)	(−2)	(−1.9)	(−1.9)	(−2)	(−2)	(−2.1)
b_2		0.098*				0.035		0.18				0.13
		(2)				(0.95)		(1.5)				(1)
b_3			−0.00024**			−0.0013***			0.000071			−0.004
			(−2.2)			(−2.8)			(0.2)			(−1.3)
b_4				0.025**		0.016*				.047		−0.0033
				(2.3)		(1.8)				(0.93)		(−0.067)
b_5					0.014**	0.011					0.025	0.037
					(2.3)	(1.6)					(1.2)	(1.4)
b_0	0.13***	0.13***	0.14***	0.1***	0.093***	0.13***	0.86***	0.88***	0.87***	0.79***	0.78***	0.86***
	(3.7)	(3.7)	(3.9)	(4.1)	(4.7)	(4.2)	(2.8)	(2.8)	(2.8)	(3)	(2.9)	(2.8)

This table reports Fama-MacBeth monthly regression of one-month individual-bond return r on lagged shorting fee (Fee), on-the-run dummy (OnRun), total size of a bond issue (IssueSize), total value of the bond issue already shorted (ValueShorted), and value of the bond issue available to lend (Inventory), separately for bonds with time to maturity no more than five years and for bonds with time to maturity longer than five years. The sample period is January 2010 to December 2012. Newey-West adjusted t-stats allowing for 12 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

Table 3
Bond return, shorting fee, and maturity

$r_{i,t} = b_1 Fee_{i,t-1} + b_2 LogMat_{i,t-1} + b_3 Fee_{i,t-1} \times LogMat_{i,t-1} + b_4 OnRun_{i,t-1} + b_5 IssueSize_{i,t-1} + b_6 ValueShorted_{i,t-1} + b_7 Inventory_{i,t-1} + b_0 + \epsilon_{i,t}$

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
b_1	−0.94	−0.98	−0.93	−1*	−0.97	−1*	−0.23	−0.23	−0.19	−0.27	−0.23	−0.25
	(−1.6)	(−1.6)	(−1.6)	(−1.7)	(−1.7)	(−1.7)	(−0.71)	(−0.73)	(−0.64)	(−0.89)	(−0.75)	(−0.97)
b_2	0.22**	0.22**	0.22**	0.21**	0.21**	0.22**	0.26**	0.26**	0.26**	0.25**	0.25**	0.26**
	(2.3)	(2.3)	(2.3)	(2.3)	(2.3)	(2.4)	(2.4)	(2.5)	(2.5)	(2.5)	(2.5)	(2.5)
b_3							−0.78**	−0.83**	−0.81**	−0.82**	−0.81**	−0.84**
							(−2.1)	(−2.1)	(−2.1)	(−2.2)	(−2.2)	(−2.1)
b_4		0.078*				0.049**		0.09*				0.055*
		(1.9)				(2.4)		(1.8)				(1.9)
b_5			0.0015**			0.001			0.0018**			0.0014*
			(2.3)			(1.5)			(2.7)			(1.7)
b_6				0.023		0.02				0.024		0.022
				(1.2)		(0.77)				(1.3)		(0.9)
b_7					0.01	−0.0033					0.011	−0.0052
					(1.4)	(−0.59)					(1.6)	(−0.8)
b_0	0.15***	0.15***	0.1**	0.14**	0.13**	0.11**	0.12**	0.12**	0.054	0.1**	0.094*	0.064*
	(2.7)	(2.7)	(2.1)	(2.4)	(2.3)	(2.3)	(2.6)	(2.5)	(1.5)	(2.1)	(1.9)	(1.9)

This table reports Fama-MacBeth monthly regression of one-month individual-bond return r on lagged shorting fee (Fee), log of time to maturity (LogMat), interaction of Fee and LogMat, on-the-run dummy (OnRun), total size of a bond issue (IssueSize), total value of the bond issue already shorted (ValueShorted), and value of the bond issue available to lend (Inventory). The sample period is January 2010 to December 2012. Newey-West adjusted t-stats allowing for 12 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

structure of bond returns on average. Columns (7) to (12) reproduce the analysis in Columns (1) to (6), but add $Fee \times LogMat$ as an extra control variable. The extra interaction term is estimated at -0.84 with a t-statistic of -2.1. The negative relationship between *Fee* and bond excess returns is thus more pronounced for long-maturity than short-maturity bonds.

Of course, this analysis may be incomplete since hedge funds can bet on inflation in other ways, such as repos and futures market, and we are only using information from the securities lending market. However, we believe the securities lending market is the most efficient to eliminate the mispricing generated by demand shocks coming from bond mutual funds. First, even though there is a growing bilateral repo market that hedge funds can use in their arbitrage strategy, the literature on repos (see, e.g., Gorton and Metrick 2012; Copeland et al. 2012; Krishnamurthy, Nagel, and Orlov 2014) find that the main motivation for bilateral repos is for institutions to engage in leveraged transactions, much in the same way as trilateral repos. The difference appears to be that whereas trilateral repos are more stringent in the collateral pool requirements, bilateral repos are more lax. This literature is in its infancy as researchers are struggling even to get an aggregate number for how big such volumes are. Nonetheless, bearish bets on inflation do not appear to be a key motivation for these transactions. Second, we collected anecdotal evidence from a practitioner, with extensive experience operating in Treasuries markets, who pointed out that hedge funds that want to short a particular maturity of Treasury, say the 30-year bond, would go directly into the securities lending market and borrow the Treasury. He also pointed out that Treasury futures and swaps were often not the most efficient way for them to short a particular Treasury bond because there is no guarantee of delivery of the particular issue that the hedge fund wanted to short in the first place.

4. Inflation Disagreement and the Yield Curve

The central prediction of our model is that when the aggregate supply of Treasuries is low, short-sale constraints are more likely to bind, and as a result the slope of the term structure of bond excess returns flattens or turns negative when there is more inflation uncertainty.

4.1 Flatness of term structure to inflation uncertainty

Before we test our central prediction, we begin by studying the relationship between the yield curve and inflation uncertainty from the perspective of the liquidity premium hypothesis, which predicts that the yield curve should steepen with inflation uncertainty. We estimate the following linear time-series regression separately for Treasuries with different maturity (k):

$$\tilde{R}_t^{(k)} = \delta_0^{(k)} + \delta_1^{(k)} \times Disagreement_{t-1} + \delta_2^{(k)} \times X_{t-1} + \epsilon_t. \tag{9}$$

$\tilde{R}_t^{(k)}$ is the realized 1-year holding period return of the k-th maturity bond in excess of the one-year bond. $Disagreement_{t-1}$ is disagreement of inflation forecasts from the Michigan Surveys lagged one month. X_{t-1} can potentially include predictor variables from the literature including CP_{t-1}, LN_{t-1} and NBER recession dates. Notice that this time-series regression is being estimated separately for each maturity k. Under the liquidity premium hypothesis, we expect the average return from holding the bond with maturity k to increase with uncertainty and that the effect of inflation uncertainty on bond risk premiums should increase with the bond's maturity so that $\delta_1^{(k)}$ should be positive and increasing with k.

Panel A of Table 4 reports the estimation of Equation (9) and shows that, on average, bond excess returns *decrease* with inflation disagreement. The first six columns present estimates of $\delta_1^{(k)}$ without controlling for the *CP* factor. At every maturity (k), $\delta_1^{(k)}$ is estimated to be negative. At the highest maturity of 10+ years, the coefficient is -2.2 with a t-statistic of 1.1. While none of these coefficients are statistically significant, the liquidity premium hypothesis stipulates that that these coefficients should be *positive*, so that these results are inconsistent with this hypothesis. To get a sense of the economic magnitudes, we consider the 10+ year bond. One standard deviation of $Disagreement_{t-1}$ is 1.5%. A one-standard deviation increase in disagreement leads to lower expected returns for the 10+ year bond of about -2.2 times 1.5% or nearly 3.3%. One standard deviation of the 10+ bond return is around 15%. So this is nearly 22% of the standard deviation of long-term bond returns, which is a sizable economic effect.

Columns (7) to (12) show the same set of estimates when the *CP* factor is included in the set of controls X_{t-1}. The estimated coefficients are similar but are now statistically significant at standard confidence levels. On long bonds (10+ year maturity), $\delta_1^{(10+)}$ is estimated at -2.9 with a t-statistic of 2.1. $\delta_1^{(5)}$ and $\delta_1^{(5-10)}$ are also negative and significant at the 1 percent confidence level. That controlling for the CP leads to more significant estimates is not surprising. To the extent that the *CP* factor is capturing time-varying risk tolerance of bond investors, controlling for the *CP* factor in our estimation allows us to isolate the pure effect of disagreement on bond returns. Additionally, controlling for the *CP* soaks up large variations in bond excess returns, which helps improve the precision of our measure of inflation uncertainty. In unreported regressions, we show that controlling for business cycle variables shown in the literature to explain bond returns, such as *LN* and *End of Recession*, does not affect these estimated effects.

In panel B of Table 4, we directly investigate the effect of inflation disagreement on the slope of the term structure of bond excess returns. To this end, we estimate the following model:

$$\widehat{Slope}_t = \nu_0 + \nu_1 \times Disagreement_{t-1} + \nu_2 \times X_{t-1} + \psi_t, \tag{10}$$

Table 4
Disagreement about Inflation and Excess Returns

Panel A. Excess returns, by maturity. $\tilde{R}_t = \delta_1 Disagreement_{t-1} + \delta_2 CP_{t-1} + \delta_0 + \epsilon_t$

	2 years	3 years	4 years	5 years	5/10 years	10+ years	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	−0.13	−0.33	−0.55	−0.78	−1.4	−2.2	−0.23	−0.5	−0.78	−1.1*	−1.8*	−2.9**
	(−0.48)	(−0.68)	(−0.84)	(−0.96)	(−1.1)	(−1.1)	(−1.1)	(−1.4)	(−1.6)	(−1.7)	(-2)	(−2.1)
δ_2							0.3***	0.53***	0.72***	0.89***	1.4***	2.2***
							(4.3)	(3.8)	(3.5)	(3.3)	(3)	(3)
δ_0	1.3	2.8	4.4	5.9*	9.6*	14*	1.7*	3.4**	5.1**	6.8**	11***	17***
	(1.2)	(1.4)	(1.6)	(1.7)	(1.9)	(1.8)	(1.9)	(2.1)	(2.4)	(2.5)	(2.8)	(2.8)
# Obs	411	411	411	411	411	411	411	411	411	411	411	411
R^2	0.011	0.021	0.03	0.038	0.048	0.048	0.21	0.2	0.2	0.2	0.21	0.22

Panel B. Slope. $\hat{Slope}_t = \nu_1 Disagreement_{t-1} + \nu_2 CP_{t-1} + \nu_3 LN_{t-1} + \nu_4 End of Recession_{t-1} + \nu_0 + \psi_t$

	ν_1	ν_2	ν_3	ν_4	ν_0
Model1	−0.18				1.1*
	(−1.2)				(1.9)
Model2	−0.23**	0.17***			1.3***
	(−2.2)	(2.7)			(2.8)
Model3	−0.17		0.19**		0.77
	(−1.4)		(2.4)		(1.3)
Model4	−0.17			−0.11	1.1*
	(−1.1)			(−0.27)	(1.8)
Model5	−0.19**	0.3***	0.09	−0.73*	0.62
	(−2.3)	(5.7)	(1.5)	(−1.8)	(1.5)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. The dependent variable in panel A is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one year bond returns. In panel B, the dependent variable is slope. CP Factor is the factor from Cochrane and Piazzesi (2005). Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

where $Slope_t$ is the estimated coefficient in the monthly cross-sectional regression of bond excess return on bond maturity in month t and X_{t-1} potentially includes similar control variables as before. As before, the liquidity premium hypothesis would imply that ν_1 is positive, that is that as inflation uncertainty increases, the slope of the term structure of bond risk premiums rise.[22]

Panel B of Table 4 shows the opposite is true. Model 1 does not include any of the additional controls X_{t-1}. ν_0, the average slope of the term structure of bond excess returns is estimated at 1.1 with a Newey-West adjusted t-statistic of 1.9, which implies that on average, the term structure of bond excess returns slopes up, although not very significantly. Importantly, ν_1 is estimated at -0.18 with a t-statistic of -1.2. The negative sign of ν_1 suggests that the term structure flattens out as disagreement increases, instead of getting steeper as implied by liquidity premium hypothesis. Note, however, that the estimated ν_1 is not statistically significant at standard confidence levels.

Model 2 includes the *CP* factor as additional control. The results are qualitatively similar to Model 1 but with a stronger statistical significance. In particular, the point estimate for ν_1 is now significantly negative, equal to -0.23 with a t-statistic of 2.2, which is again inconsistent with the liquidity premium hypothesis. In contrast, Model 2 shows that a higher *CP* factor leads to a steeper term structure as ν_2 is estimated to be significantly positive. Models 3 and 4 include alternatively the *LN* factor and the *End of Recessions* dummies as control variables in Equation (10). Finally, Model 5 use both *CP* and *LN* as control variables. Across these 5 specifications, the point estimate for ν_1 hovers around 0.2 and is statistically significant at the 5% confidence level in Model 2 and 5. We conclude from this analysis that the liquidity premium hypothesis – the hypothesis that inflation uncertainty leads to an increase in the slope of the term structure of bond excess returns—does not find much support in the data. We next explore the role of aggregate Treasury supply in explaining the relationship between the term structure of bond returns and inflation disagreement.

4.2 The role of aggregate supply

To see if the failure of the liquidity premium hypothesis is due to the speculative forces we have outlined in our model, we test the model's central prediction: the effect of disagreement on the slope of the term structure of bond returns should be much larger when the aggregate (maturity-weighted) supply of bonds

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[22] Because the slope estimate is a left-hand-side variable instead of a right-hand-side variable in the second stage regression, there is no need for a correction of the errors-in-variables problem for the standard errors (see, e.g., Shanken 1992).

Table 5
Disagreement about inflation, bond supply, and excess returns

Panel A. Excess returns, by maturity.
$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	0.8***	1.5***	2***	2.5***	3.6***	5.1***
	(4.6)	(4.6)	(4.5)	(4.4)	(3.8)	(3.3)
δ_2	−1.6***	−3***	−4.3***	−5.4***	−7.8***	−11***
	(−3.7)	(−3.9)	(−4)	(−4)	(−3.8)	(−3.5)
δ_3	−3.1***	−5.7***	−8***	−9.8***	−14***	−19***
	(−3.9)	(−4)	(−3.9)	(−3.8)	(−3.4)	(−2.9)
δ_0	6.8***	13***	19***	24***	34***	46***
	(3.1)	(3.3)	(3.4)	(3.5)	(3.4)	(3.1)
# Obs	411	411	411	411	411	411
R^2	0.2	0.21	0.21	0.21	0.2	0.19

Panel B. Slope.
$\widehat{Slope}_t = \nu_1 Disagreement_{t-1} \times Supply_{t-1} + \nu_2 Disagreement_{t-1} + \nu_3 Supply_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_0
0.36***	−0.79***	−1.3**	3.1***
(2.8)	(−3)	(−2.4)	(2.7)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). In panel A, the dependent variable is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one year bond returns. In panel B, the dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

is low. To do so, we simply estimate the following equation:

$$\begin{aligned}\tilde{R}_t^{(k)} = \delta_0^{(k)} + \delta_1^{(k)} \times Disagreement_{t-1} \times Supply_{t-1} \\ + \delta_2^{(k)} \times Disagreement_{t-1} + \delta_3^{(k)} \times Supply_{t-1} \\ + \delta_4^{(k)} X_{t-1} + \delta_5^{(k)} X_{t-1} \times Supply_{t-1} + \epsilon_{k,t},\end{aligned} \tag{11}$$

where $\tilde{R}_t^{(k)}$ is the excess 1-year holding period return of a bond with maturity k, $Supply_{t-1}$ is the maturity-weighted Treasury supply variable from Greenwood and Vayanos (2014) and X_{t-1} is potentially the same set of control variables as in Equation (10). This specification allows the effect of inflation disagreement on bond returns to vary with the Treasury supply, which is a key feature of our model with investors disagreement and short-sale restriction. Precisely, our theory predicts that $\delta_1^{(k)}$ should be positive, that is, the effect of inflation disagreement on bond returns should decrease as supply rises.

Panel A of Table 5 presents the results from the estimation of Equation (11) when no control variables are added to the equation. As predicted by our theory, $\delta_1^{(k)}$ is positive and highly significant across all maturities. For instance, $\delta_1^{(10+)}$ is estimated at 5.1 with a t-statistic of 3.3. At the 90th percentile of the

Supply distribution (i.e., 4.29), a one standard deviation increase in inflation disagreement (about 1.5) leads to an increase in the excess returns of 10+ year bonds of about $\delta_1^{(10+)} \times 1.5 \times 4.29 + \delta_2^{(10+)} \times 1.5$ or about 16 percent. At the 10th percentile of the Supply distribution (1.72), the same increase in inflation disagreement leads to a *decrease* in the excess returns of 10+ year bonds of $\delta_1^{(10+)} \times 1.5 \times 1.72 + \delta_2^{(10+)} \times 1.5$ or about 3.3 percent.

We also estimate $\delta_3^{(k)}$ to be negative and significant across bonds with different maturities. Quantitatively, a one standard deviation increase in $Supply_{t-1}$ is around 1. At the median inflation disagreement level (4.1), a one standard deviation increase in supply leads to a negligible increase in the 10+year excess returns of about $\delta_1^{(10+)} \times 4.1 \times 1 + \delta_3^{(10+)} \times 1 = 1.9$ percent.[23]

We then analyze directly the effect of inflation disagreement on the term structure of bond excess returns depends on bond supply. To do so, we estimate the following equation:

$$\begin{aligned} \hat{Slope}_t = \nu_0 + \nu_1 \times Disagreement_{t-1} \times Supply_{t-1} \\ + \nu_2 \times Disagreement_{t-1} + \nu_3 Supply_{t-1} \\ + \nu_4 \times X_{t-1} + \nu_5 \times X_{t-1} \times Supply_{t-1} + \psi_t, \end{aligned} \tag{12}$$

Our theory predicts that because long bonds are more exposed to inflation disagreement, the negative effect of disagreement on bond returns should significantly decrease with the bond's maturity, especially if Supply is low enough. In other words, our model predicts that ν_1 should be positive

Panel B of Table 5 reports the estimation of Equation (12) when no control variables are added in the estimation. As predicted by our theory, ν_1 is estimated to be positive at 0.36 with a t-statistic of 2.8. To get a better intuition for these findings, we turn to Figure 5. Figure 5 plots the term structure of bond returns for high and low supply periods (defined as months in the top vs. bottom tercile of the supply distribution) separately for months with high and low inflation disagreement (defined as the top vs. bottom tercile of the disagreement distribution). Figure 5 shows that whenever disagreement is limited, the term structure of bond excess returns does not depend on Treasury supply. However, when inflation disagreement is large, a low supply leads to a downward sloping term structure of bond returns, while it remains upward sloping whenever supply is low. We interpret this downward-sloping term structure when aggregate supply is low and disagreement is high as the result of the inflation-betting channel we emphasize in our model.

In Table 6, we replicate the regressions of Table 5 but add the *CP* factor as an additional control X_{t-1}. The point estimates for $\delta_1^{(k)}$ and $\delta_2^{(k)}$ are generally very much in line with the estimation of Table 5, but they have stronger statistical

[23] While Greenwood and Vayanos (2014) reports that an increase in supply leads to a significant reduction in bond excess returns, we find that over our more recent and shorter sample period, the effect is closer to 0.

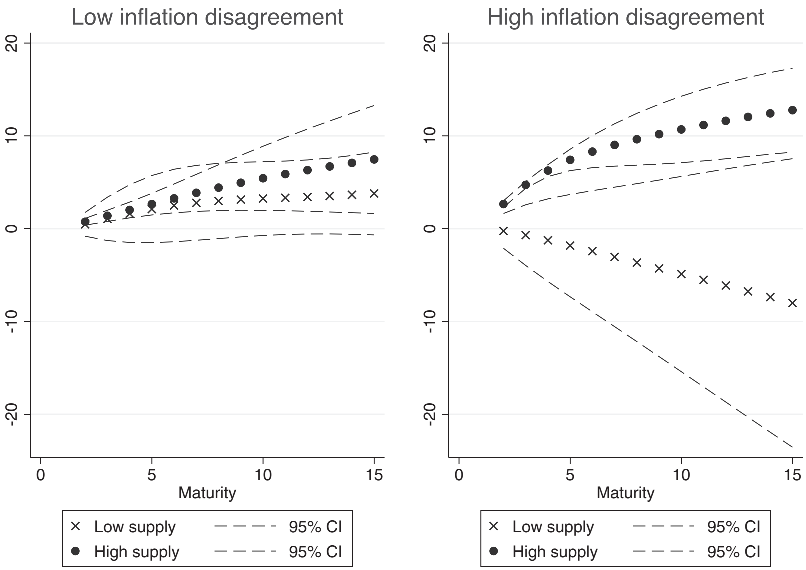

Figure 5
Bond excess returns for low versus high aggregate supply months conditional on inflation disagreement
The left panel plots the average bond 1-year returns in excess of the 1-year Treasury bill for top and bottom tercile aggregate supply months when inflation disagreement is in the bottom tercile of disagreement over the period of 1978–2012. The right panel is similar, except that it is for months when inflation disagreement is in the top tercile over the period of 1978–2012. The dashed lines correspond to 95% confidence interval.

significance. For instance, in Table 6, $\delta_1^{(10+)}$ is estimated at 4.1, with a t-statistic of 3.6, compared to 5.1 with a t-statistic of 3.3 in Table 5. Similarly, $\delta_2^{(10+)}$ is -10 with a t-statistic of 4.8 in Table 6 compared with -11 and a t-statistic of 3.5 in Table 5.

Panel B of Table 6 is also very similar with the results obtained in Table 5: when controlling for the CP factor, ν_1 is estimated at 0.28 with a t-statistic of 2.9, which can be compared with an estimate for ν_1 without controls of 0.36 with a t-statistic of 2.8. The inclusion of CP_{t-1} does not affect the conclusions drawn from Table 5.

Ludvigson and Ng (2009) show that business cycle variables are useful in predicting bond excess returns. As evident on Figure 1, the dispersion of inflation forecast is counter-cyclical: this time series spikes during all the recessions in the sample. Hence, it is possible that some of the predictive power that we document is associated with countercyclical risk premiums, rising during recessions, rather than the time-varying effects of short-sale constraints. In Table 7, we replicate the regressions of Table 5 but add now the *LN* factor from Ludvigson and Ng (2009) as an additional control X_{t-1}. In Table 8, we control instead for *End of Recession* when estimating Equations (11) and (10), which is a dummy variable equal to 1 for the last 3 months of NBER recessions periods and zero otherwise. There is widespread belief that the yield curve

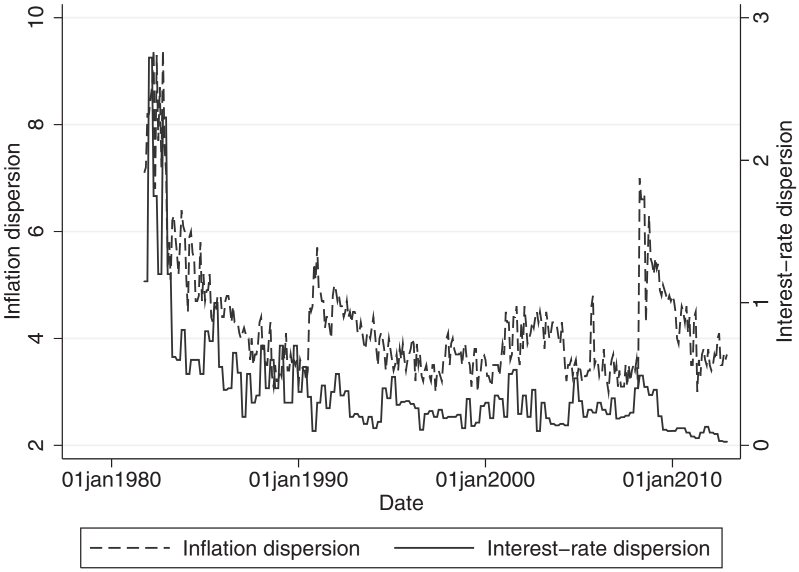

Figure 6
Time series of inflation disagreement and interest-rate disagreement
This figure plots the interquartile range in monthly inflation forecasts from Michigan Survey and Treasury-bill rate forecasts from Survey of Professional Forecasters.

inverts before recessions and steepens at the end of recessions. However, our analysis in Table 8 does not find support for this belief. More importantly, our estimates of both $\delta_1^{(k)}$ and $\delta_2^{(k)}$ in Tables 7 and 8 are virtually identical to those obtained in Table 5. Overall, the analysis in this section strongly supports the view that in a low supply environment, an increase in inflation uncertainty leads to a flatter term structure of bond expected returns, which is the main prediction from our theoretical analysis.[24]

4.3 Subperiod Analysis

Our model implicitly assumes that inflation risk increases bond risk premiums. Yet, if inflation is procyclical, as it appears to have been since the late 1990s, higher inflation makes bonds more of a hedge, and one would expect higher inflation risk to reduce expected bond returns and yields (Burkhardt and Hasseltoft 2012). Similarly, David and Veronesi (2013) find support for the notion that bond volatility and bond yields were positively correlated in the late 1970s and early 1980s, because of an increase in the beliefs of moving to a high-inflation regime. Conversely, starting in the late 1990s, the relation between yields and bond volatility becomes negative because of uncertainty on whether a deflationary regime may occur.

[24] In Appendix Table A7, we also control for more recent term structure variables from the literature.

Table 6
Disagreement about inflation, bond supply, and excess returns: Controlling for CP factor

Panel A. Excess returns, by maturity.
$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_4 CP_{t-1} \times Supply_{t-1} + \delta_5 CP_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	0.64***	1.2***	1.7***	2.1***	2.9***	4.1***
	(4.5)	(4.6)	(4.6)	(4.4)	(4)	(3.6)
δ_2	−1.4***	−2.8***	−4***	−5***	−7.4***	−10***
	(−4.1)	(−4.5)	(−4.6)	(−4.7)	(−4.8)	(−4.8)
δ_3	−2.6***	−5***	−7***	−8.6***	−12***	−16***
	(−3.9)	(-4)	(-4)	(−3.9)	(−3.6)	(−3.2)
δ_4	−0.077	−0.14	−0.2	−0.27	−0.52	−0.99*
	(−1.5)	(−1.4)	(−1.4)	(−1.4)	(−1.5)	(−1.8)
δ_5	0.45***	0.79**	1.1**	1.4**	2.5***	4.4***
	(2.6)	(2.5)	(2.5)	(2.5)	(2.7)	(2.9)
δ_0	7***	14***	19***	24***	35***	48***
	(3.4)	(3.7)	(3.8)	(4)	(4.1)	(4.1)
# Obs	411	411	411	411	411	411
R^2	0.32	0.31	0.31	0.31	0.3	0.31

Panel B. Slope.
$\widehat{Slope}_t = \nu_1 Disagreement_{t-1} \times Supply_{t-1} + \nu_2 Disagreement_{t-1} + \nu_3 Supply_{t-1} + \nu_4 CP_{t-1} \times Supply_{t-1} + \nu_5 CP_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_4	ν_5	ν_0
0.28***	−0.75***	−1.1***	−0.086*	0.37***	3.3***
(2.9)	(−4.2)	(−2.6)	(−1.8)	(3)	(3.7)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978-2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). CP Factor is the factor from Cochrane and Piazzesi (2005). In panel A, the dependent variable is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one year bond returns. In panel B, the dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

In this context, it could be that inflation uncertainty as proxied by dispersions of Michigan forecasts might be negatively related to future bond returns because of this hedging mechanism as opposed to our proposed mechanism highlighting disagreement and short-sales constraints. However, given the nature of inflation risk historically, this alternative inflation-hedge mechanism is most likely to be prominent after the late 1990s. In Table 9, we reproduce the analysis of Table 5 for different subperiods. Columns (1) and (2) exclude 1978–1979, which corresponds to the hyperinflation regime in the United States before the "Volcker" interventions. $\delta_1^{(2)}$ is estimated at 0.7 with a t-statistics of 3 and $\delta_1^{(10+)}$ is estimated at 4.1 with a t-statistic of 2.3. Both estimates are very similar in magnitudes to those obtained in panel A of Table 5. As reported on the last raw of Table 5, panel A, the p-value for the difference of the point estimates using the sample of 2-year bonds versus 10+ year bonds is 0.12. The fact that this difference is only marginally significant is not surprising since the late

Table 7
Disagreement about inflation, bond supply, and excess returns: Controlling for LN factor

Panel A. Excess returns, by maturity.

$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_4 LN_{t-1} \times Supply_{t-1} + \delta_5 LN_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	0.68***	1.3***	1.8***	2.2***	3.2**	4.8**
	(2.6)	(2.6)	(2.6)	(2.6)	(2.4)	(2.3)
δ_2	−1.5**	−2.8**	−4***	−5***	−7.4**	−10**
	(−2.4)	(−2.5)	(−2.6)	(−2.6)	(−2.5)	(−2.3)
δ_3	−2.7**	−5**	−7**	−8.6**	−12*	−16*
	(−2.1)	(−2.2)	(−2.2)	(−2.2)	(−2)	(−1.7)
δ_4	−0.17*	−0.29*	−0.41*	−0.53*	−0.98**	−1.9**
	(−1.8)	(−1.7)	(−1.7)	(−1.8)	(−2)	(−2.4)
δ_5	0.78**	1.4**	1.8**	2.3**	3.7**	6.6***
	(2.6)	(2.5)	(2.5)	(2.4)	(2.5)	(2.7)
δ_0	6.4*	12*	18**	22**	30*	38
	(1.8)	(1.9)	(2)	(2)	(1.8)	(1.5)
# Obs	310	310	310	310	310	310
R^2	0.34	0.33	0.32	0.31	0.29	0.28

Panel B. Slope.

$\hat{Slope}_t = \nu_1 Disagreement_{t-1} \times Supply_{t-1} + \nu_2 Disagreement_{t-1} + \nu_3 Supply_{t-1} + \nu_4 LN_{t-1} \times Supply_{t-1} + \nu_5 LN_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_4	ν_5	ν_0
0.35**	−0.74**	−1.1	−0.17**	0.54***	2.3
(2.1)	(−2.1)	(−1.4)	(−2.6)	(2.7)	(1.2)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). LN is the bond return factor $F6$ in Ludvigson and Ng (2009). In panel A, the dependent variable is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one year bond returns. In panel B, the dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

seventies is important for the identification of our main effects. However, the fact that the point estimates remains similar despite the exclusion of this period comforts us in the robustness of the estimated effect.

Columns (3) and (4) exclude 1980–1984, which corresponds to the interventions of the Federal Reserve Bank to bring down inflation and associated recession. Columns (5) and (6) drop the 1985–1989 period, Columns (7) and (8) exclude 1990–1994, Columns (9) and (10) 1995–1999, and Columns (11) and (12) drop 2000–2004. For all these subperiod, all the estimated $\delta_1^{(2)}$ and $\delta_1^{(10+)}$ are statistically significant at the 1% confidence level and close to their estimated value on the full sample. In all these subsamples, the difference between these point estimates is always significant at least at the 6% confidence level. Finally, Columns (13) and (14) exclude 2005-2012, a period during which quantitative easing makes it hard to estimate what fraction of Treasury supply

Table 8
Disagreement about inflation, bond supply, and excess returns: Controlling for recessions

Panel A. Excess returns, by maturity.

$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_4 End\ of\ Recession_{t-1} \times Supply_{t-1} + \delta_5 End\ of\ Recession_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	0.77***	1.4***	2***	2.5***	3.7***	5.3***
	(4.4)	(4.4)	(4.4)	(4.3)	(3.8)	(3.3)
δ_2	−1.5***	−2.9***	−4.2***	−5.3***	−7.8***	−11***
	(−3.7)	(−4)	(−4.1)	(−4.1)	(−3.9)	(−3.4)
δ_3	−3***	−5.6***	−7.9***	−9.8***	−14***	−20***
	(−3.9)	(−3.9)	(−3.9)	(−3.8)	(−3.4)	(−3)
δ_4	0.84	1.2	1.2	1	0.15	−1.8
	(1.2)	(0.99)	(0.75)	(0.54)	(0.057)	(−0.5)
δ_5	−2.9	−4.1	−4.6	−4.7	−4	−0.61
	(−1.4)	(−1.2)	(−1)	(−0.85)	(−0.5)	(−0.054)
δ_0	6.5***	13***	19***	23***	34***	46***
	(3.1)	(3.3)	(3.5)	(3.5)	(3.5)	(3.1)
# Obs	411	411	411	411	411	411
R^2	0.21	0.21	0.22	0.22	0.21	0.19

Panel B. Slope.

$\hat{Slope}_t = \nu_1 Disagreement_{t-1} \times Supply_{t-1} + \nu_2 Disagreement_{t-1} + \nu_3 Supply_{t-1} + \nu_4 End\ of\ Recession_{t-1} \times Supply_{t-1} + \nu_5 End\ of\ Recession_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_4	ν_5	ν_0
0.38***	−0.82***	−1.4**	−0.31	0.39	3.2***
(2.8)	(−3)	(−2.5)	(−1.1)	(0.44)	(2.7)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). *End of Recession* is a dummy variable equal to 1 for the last 3 months of the NBER recessions period. In panel A, the dependent variable is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one year bond returns. In panel B, the dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

is held by the public.[25] Again, the coefficient estimates are virtually similar on this subperiod.

4.4 Robustness checks

Having established our baseline results, we turn to a series of robustness checks. For the sake of brevity, these tables are provided in the Internet Appendix.

4.4.1 Discretizing disagreement and supply measures. In Table A1, we consider two specification checks to investigate whether outliers in $Disagreement_{t-1}$ and $Supply_{t-1}$ drive our results. In panel A, we replace our continuous $Disagreement_{t-1}$ measure with dummies for disagreement terciles (High, Medium and Low) and re-estimate Equation (11). Table A1 shows that

[25] In the Online Appendix, we also include the supply of repos in the market into Treasury supply.

Table 9
Disagreement about inflation, bond supply, and excess returns: Subperiods

$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	10+ years	2 years	10+ years	2 years	10+ years	2 years	10+ years	2 years	10+ years	2 years	10+ years	2 years	10+ years
	≥ 80		$\notin [80,84]$		$\notin [85,89]$		$\notin [90,94]$		$\notin [95,99]$		$\notin [00,04]$		$\notin [05,12]$	
δ_1	0.7***	4.1**	0.62***	4.5***	0.8***	4.8***	0.97***	6.7***	0.84***	5.6***	0.83***	5.1***	0.79***	5.4***
	(3)	(2.3)	(4.8)	(3.3)	(4.4)	(3.3)	(4.2)	(3.2)	(4.3)	(3.3)	(4.5)	(3.1)	(3.9)	(3)
δ_2	−1.4**	−8.9**	−1.4***	−12***	−1.6***	−10***	−1.8***	−13***	−1.7***	−12***	−1.6***	−11***	−1.5***	−12***
	(−2.2)	(−2)	(−4.6)	(−4.6)	(−3.7)	(−3.5)	(−4)	(−3.7)	(−3.8)	(−3.6)	(−3.5)	(−3.1)	(−2.9)	(−3.1)
δ_3	−2.8***	−16**	−2.3***	−17***	−3***	−17***	−3.6***	−24***	−3.3***	−21***	−3.2***	−18***	−2.9***	−21**
	(−2.8)	(−2.1)	(−3.8)	(−2.9)	(−3.7)	(−2.8)	(−4)	(−3)	(−3.7)	(−2.8)	(−3.7)	(−2.6)	(−2.9)	(−2.6)
δ_0	6.6**	40**	6.1***	48***	6.7***	41***	7.5***	51***	7.3***	49***	6.8***	41**	5.9**	49**
	(2.3)	(2.1)	(3.2)	(4)	(3)	(3.1)	(3.4)	(3.4)	(3.2)	(3.2)	(2.8)	(2.5)	(2)	(2.5)
# Obs	389	389	351	351	351	351	351	351	351	351	351	351	334	334
R^2	0.14	0.089	0.19	0.2	0.22	0.22	0.22	0.23	0.21	0.21	0.24	0.19	0.22	0.19
p-value for difference	0.12		0.03**		0.053*		0.031**		0.039**		0.042**		0.023**	

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply, excluding various subperiods. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). The dependent variable is the 1-year holding period return of 2- (Columns 1, 3, 5, 7, 9, 11, and 13) and 10+ year (Columns 2, 4, 6, 8, 10, 12, and 14) bonds in excess of the one year bond returns. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively. Column 1 and 2 exclude 1978-1979, Columns 3 and 4 exclude 1980-1984, Columns 5 and 6 exclude 1985–1989, Columns 7 and 8 exclude 1990–1994, Columns 9 and 10 exclude 1995–1999, Columns 11 and 12 exclude 2000-2004, and Columns 13 and 14 exclude 2005–2012. The line p-value of difference is the p-value of the difference of the point estimates of the 2-year return and the 10+ year return regression.

the effect of disagreement on bond returns in a low supply environment is mostly coming from High disagreement months. It also shows that this effect is not driven by outliers in the distribution of inflation disagreement. In panel B, we perform an analogous exercise but discretize the supply measure ($Supply_{t-1}$) and find that this does not affect our previous conclusion that aggregate supply is important in moderating the effect of disagreement on bond returns.

4.4.2 Using Fama bonds file to calculate bond returns. Table A2 reproduces Table 5 using Fama Bonds instead of the FED Bond Pricing data. In panel A, across all maturities, we find positive and significant estimates for $\delta_1^{(k)}$ and negative and significant estimates for $\delta_2^{(k)}$. In panel B, we report a positive estimate for ν_1 and significant at the 5% confidence level, which again shows that inflation disagreement leads to a flatter term structure of Fama bond returns mostly when Treasury supply is low. Overall, the point estimates from Table A2 are smaller in magnitudes and marginally less significant than the results obtained in Table 5, but nonetheless support the paper's main hypothesis.

4.5 Using bond yields instead of bond holding period returns

Table A3 reproduces the specifications of Table 5, but uses the yield spread as a dependent variable. Panel A shows that, at all horizon, the negative effect of disagreement on the yield spread is attenuated whenever supply is high, which is consistent with our results on bond returns shown in panel A of Table 5. Panel B shows that the yield curve flattens out when disagreement is high and Treasury supply is low. These effects are all significant at the 1% confidence level. To account for the small sample property of this estimation, as the well as the persistence in the yields, we also report bootstrapped p-values for the estimation in panel B. Following Cochrane and Piazzesi (2005) and Ludvigson and Ng (2009), we run an AR(12) regression of slope and each of the right-hand-side variables, and bootstrap the residuals. We then report the p-value of the actual Newey-West t-statistic relative to the empirical distribution of the Newey-West t-statistic from 100,000 simulations.

4.5.1 Using Livingston survey to measure disagreement about inflation expectation. Table A4 reproduces the analysis of Table 4, but constructs the inflation disagreement measure from the Livingston Survey instead of the Michigan Survey. The Livingston Survey is available from 1952 to 2012, but the survey only samples semiannually, in the months of June and December. The correlation between the two survey measures is high, at around 0.8. While the Livingston Survey provides far less variations in $Disagreement_{t-1}$ than the Michigan survey (Figure A1), the Livingston series might better capture the expectations of professional investors.

Panel A of Table A4 shows results very similar to those estimated on panel A of Table 4. $\delta_1^{(10+)}$ is estimated at -6.5 with a t-statistic of 2. As in Table 4, ν_1

is consistently estimated negative across the different estimation models used and is statistically significant in three out of the 5 specifications. Note that the point estimates across Table A4 and Table 4 cannot be directly compared since the standard deviation of disagreement is three times larger for the Michigan sample. Nonetheless, when adjusting for the different standard deviations, we do obtain very similar magnitudes for the coefficients of interest. Finally, Figure A2 shows the term structure of bond excess returns for high and low disagreement months when disagreement is measured using the Livingston Survey. The figure is essentially similar to Figure 2.

4.6 Alternative methods of calculating standard errors

We present a number of alternative methods for calculating standard errors for our main results, namely, the estimation of ν_1 in panel A of Table 5. In panel A of Table A5, we present Hansen-Hodrick t-statistics, as in Hansen and Hodrick (1980), to address finite sample issues of the Newey and West (1987) t-statistics. In also these specifications, we also follow Cochrane and Piazzesi (2005) and Ludvigson and Ng (2009), and compute bootstrap p-values. To do so, we run an AR(12) regression of slope and each of the right-hand-side variables, and bootstrap the residuals. Panel A shows the bootstrap p-value of the actual Newey-West t-statistic relative to the empirical distribution of the Newey-West t-statistic from 100,000 simulations. The significance of the estimated effects is very similar using these alternative estimation for standard errors.

We also explore the statistical significance of our result using a different estimation technique. In panel C, we run a regression pooling bond of all maturities, where the excess 1 year-holding returns of a maturity-(k) bond is regressed against inflation disagreement in the previous month, maturity and the interaction of these two variables. This specification is akin to the one used in the context of our two-stage procedure in panel B of Table 4. Panel B considers a similar pooled regression but include the triple interactions of maturity, inflation disagreement and Treasury supply. For these two estimations, we show t-statistics computed using Newey-West allowing for thirteen lags, Hansen-Hodrick t-statistics, standard errors clustered at the monthly level. We also show the results using time fixed effects, as well as time and maturity fixed effects.

Our main conclusions remain unaffected by these various robustness checks.

5. Additional Analysis

We conclude the paper by extending our analysis in a number of key dimensions: (1) attempting to disentangle disagreement from inflation risk, (2) verifying that inflation disagreement is AR(1), (3) measuring the effect of interest rate disagreement, and (4) relating our aggregate disagreement measure to trading volume and spreads.

5.1 Disentangling disagreement and inflation risk

Disagreement in inflation forecasts is viewed as the best measure of inflation uncertainty or risk (Giordani and Söderlind 2003) and much better than standard time-series models of inflation uncertainty such as AR processes. We thus view the estimated effect of $Disagreement_{t-1}$ as capturing the net of inflation disagreement and inflation risk.

To disentangle disagreement from inflation risk, we exploit the fact that professional forecasters in the SPF provide a standard error around their forecast.[26] This standard error can be interpreted as inflation uncertainty as assessed by profession forecasters. We first compute for each forecaster this standard deviation following Equation (3) in Giordani and Söderlind (2003) and then use the median standard error as our monthly measure of inflation risk ($InfRisk_{t-1}$). One potential issue with this measure is that forecasters tend to be overconfident and thus under-report the true standard errors (Giordani and Söderlind 2003).

Table 10 reproduces the analysis of Table 5, but includes $InfRisk_{t-1}$ as an additional control. The estimated $\delta_1^{(k)}$ and $\delta_2^{(k)}$ remain almost similar to the estimations from Table 5. In contrast, the interaction of our inflation risk measure and Treasury supply does not forecast in a significant way bond excess returns. A potential explanation for this insignificant result is that the standard errors from the SPF are noisily measured because of forecasters' overconfidence. Another interpretation is that times of high inflation risk are so tied to disagreement and the disagreement effect is strong enough that it dominates the inflation risk effect.

5.2 Term structure of disagreement

Our model, which features an AR(1) process for inflation, implicitly generates an AR(1) structure for inflation disagreement. In panel A Table 11, we verify this is the case using the SPF inflation data. In any given quarter t, the SPF provides us with data on forecasts out one quarter (Q1), two quarters (Q2), three quarters (Q3), one-year ahead (Y1) and ten-year ahead (Y1/10). We can therefore measure the term structure of dispersion in any given quarter. Table 11 shows there is more dispersion of forecasts in the near term and that this dispersion is monotonically declining as maturity increases, consistent with an AR(1) process for inflation and inflation forecast dispersion.

However, when we conduct a formal statistical test by examining the difference between the Y1 and Y1/10 forecasts or the Q1 and Y1/10 forecasts, we see that these differences are not statistically different from zero. In other words, the AR(1) process is very persistent. These findings are similar to the persistence of inflation disagreement found in Andrade et al. (2014) and more generally to studies of inflation shocks reporting that inflation follows close to a random walk (Atkeson and Ohanian 2001).

26 The Michigan household series does not provide such standard errors.

Table 10
Disagreement about inflation, bond supply, and excess returns: Controlling for inflation risk

Panel A. Excess returns, by maturity.

$\tilde{R}_t = \delta_1 Disagreement_{t-1} \times Supply_{t-1} + \delta_2 Disagreement_{t-1} + \delta_3 Supply_{t-1} + \delta_4 InfRisk_{t-1} \times Supply_{t-1} + \delta_5 InfRisk_{t-1} + \delta_0 + \epsilon_{k,t}$

	2 years	3 years	4 years	5 years	5/10 years	10+ years
δ_1	0.75***	1.4***	1.9***	2.4***	3.5***	5***
	(4.2)	(4.2)	(4.1)	(4)	(3.5)	(3.1)
δ_2	−1.5***	−2.9***	−4.1***	−5.1***	−7.6***	−11***
	(−3.3)	(−3.4)	(−3.5)	(−3.5)	(−3.4)	(−3.2)
δ_3	−3.3***	−6.2***	−8.7***	−11***	−15***	−19***
	(−4.2)	(−4.3)	(−4.2)	(−4)	(−3.4)	(−2.7)
δ_4	0.53	1.2	1.7	2	2	1.4
	(0.68)	(0.82)	(0.79)	(0.71)	(0.44)	(0.18)
δ_5	−0.57	−2.3	−3.6	−4.4	−4.4	−2.2
	(−0.2)	(−0.42)	(−0.46)	(−0.43)	(−0.27)	(−0.083)
δ_0	7.1***	14***	21***	26***	36***	47**
	(2.9)	(3.2)	(3.3)	(3.3)	(3)	(2.5)
# Obs	411	411	411	411	411	411
R^2	0.22	0.22	0.22	0.22	0.21	0.19

Panel B. Slope.

$\hat{Slope}_t = \nu_1 Disagreement_{t-1} \times Supply_{t-1} + \nu_2 Disagreement_{t-1} + \nu_3 Supply_{t-1} + \nu_4 InfRisk_{t-1} \times Supply_{t-1} + \nu_5 InfRisk_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_4	ν_5	ν_0
0.36***	−0.79***	−1.3**	−0.0096	0.08	3.1**
(2.7)	(−2.9)	(−2.1)	(−0.016)	(0.038)	(2)

This table reports linear regressions of 1-year holding excess bond returns on past-month disagreement about inflation and maturity-weighted supply. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Data on disagreement about inflation come from the Michigan Survey. They are measured as the monthly interquartile range of consumers forecast for the next year's inflation rate. The sample period is 1978–2012. Maturity-weighted supply is the maturity-weighted-debt-to-GDP ratio defined in Greenwood and Vayanos (2014). *InfRisk* is the inflation risk from Survey of Professional Forecasters. We compute standard deviation using individual forecasters forecast distribution. InfRisk is the median standard deviation across forecasters. In panel A, the dependent variable is the 1-year holding period return of 2, 3, 4, 5, 5- to 10-, and 10+ year bonds in excess of the one-year bond returns. In panel B, the dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

5.3 Trading volume as proxy for disagreement

We also connect the findings in this paper to the earlier literature on disagreement and turnover in equity market. Disagreement models typically lead to greater trading volume or turnover in a dynamic setting. Empirical research on equities often equate disagreement to turnover. One key difference in the Treasury market is that bond investors such as insurance companies might have longer investment horizons and less speculative motive than retail investors do on the equity market. As a result, there might be less speculation in the Treasury market and thus a lower effect of trading volume on returns.

Panel B, Table 11, reproduces the results in panel B of Table 4 but replace our Treasury supply measure with a measure of turnover in the Treasury market. In this regression, we are careful to control for the bid-ask SPREAD in Treasury markets so as to not capture liquidity effects of trading volume as opposed to disagreement. We see that the coefficient in front of turnover is negative and

Table 11
Additional analysis

Panel A. Term structure of disagreement

	Q1	Q2	Q3	Q4	Y1	Y1/10	Y1-Y1/10	Q1-Y1/Y10
Dispersion	0.77	0.7	0.69	0.68	0.6	0.53	0.075	0.24
	(6.7)	(8.9)	(12)	(15)	(12)	(10)	(0.91)	(1.7)

Panel B. Turnover and excess returns: Controlling for spread and CP.
$\hat{Slope}_t = \nu_1 TURNOVER_{t-1} + \nu_2 Spread_{t-1} + \nu_3 CP_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_0
−0.53	11	0.15**	1***
(−1.5)	(1)	(2.2)	(3.8)

Panel C. Disagreement and excess returns: Controlling for spread and CP.
$\hat{Slope}_t = \nu_1 Disagreement_{t-1} + \nu_2 Spread_{t-1} + \nu_3 CP_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_0
−0.31***	9.5**	0.21***	1.8***
(−3.6)	(2.2)	(3.8)	(4.7)

Panel D. Interest-rate disagreement and excess returns.
$\hat{Slope}_t = \nu_1 RateDis_{t-1} + \nu_2 CP_{t-1} + \nu_3 LN_{t-1} + \nu_4 End\ of\ Recession_{t-1} + \nu_0 + \psi_t$

ν_1	ν_2	ν_3	ν_4	ν_0
−0.13	0.27***	0.049	−0.79	0.0082
(−0.28)	(3.7)	(0.71)	(−1.6)	(0.038)

Panel A reports the forecast dispersion by forecast horizon. Q1 to Q4 are dispersions over inflation in the 1st to 4th quarter ahead. Y1 is the dispersion over inflation up to December of next year. Y1/10 is the dispersion over inflation in the next ten years. The forecasts are annualized. The quarterly data are from the Survey of Professional Forecasters from 1991Q4 to 2012. Newey-West adjusted t-stats with 40 quarterly lags are in parentheses. Panel B reports linear regressions of 1-year holding excess bond returns on past-month bond market TURNOVER. The sample period of TURNOVER is 1991–2012. The dependent variable is slope. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. Panel C reports linear regressions of 1-year holding excess bond returns on past-month inflation disagreement, controlling for past-month bond market bid-ask spread. The dependent variable is slope. The sample period is 1978–2012. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. Panel D reports linear regressions of 1-year holding excess bond returns on past-month interest-rate disagreement. *RateDis* is interquartile disagreement about Treasury-bill rate from the Survey of Professional Forecasters. The dependent variable is Slope. CP Factor is the factor from Cochrane and Piazzesi (2005). Newey-West adjusted t-stats allowing for 13 lags are in parentheses. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

the t-statistic is -1.5. In panel C, Table 11, we rerun our main analysis while controlling for spread. We find similar effects of inflation disagreement on the term structure of bond excess returns.

5.4 Interest rate disagreement

Term structure models of interest rates typically feature two sources of risk, inflation risk and also interest rate risk. We can extend our model to feature interest rate risk. Our predictions for inflation disagreement then extend over to interest rate disagreement.

In Figure 6, we plot both the inflation disagreement series from Michigan and the interest rate disagreement series from the SPF. After the hyperinflation of the late 1970s and early 1980s, there is a pronounced decline in interest rate disagreement, which then hovers around a low level (of about 50 bps). We reproduce the analysis of panel B Table 4, but replace inflation disagreement by interest rate disagreement as the main explanatory variables. The results, in panel D Table 11, show that interest rate disagreement negatively forecasts the slope of the term structure of bond excess returns, but this result is not significantly different from 0.[27]

This is puzzling to the extent that interest rate risk is considered an important factor in term structure models. We speculate on a potential rationale but leave a formal analysis for future research. If many forecasters believe in the Taylor Rule, then interest rate disagreement should simply combine both inflation disagreement and unemployment disagreement. Dräger and Lamla (2015) show that accounting for such a Taylor Rule relationship is important for understanding interest rate disagreement. However, empirically, unemployment disagreement does not forecast bond returns.[28] As a result, this might explain why interest rate disagreement does not have any additional forecasting power for predicting bond returns beyond what is contained in inflation disagreement. We believe this is an interesting topic for future research.

5.5 Lending fees and time-varying bond returns

In this final subsection, we want to connect our findings in Table 3, whereby high lending-fee bonds underperform low lending-fee bonds, to inflation disagreement. To the extent that inflation disagreement is an aggregate indicator of speculative demand, we hypothesize that this underperformance relationship documented in Table 3 is more prominent when inflation disagreement is high.

We conduct this analysis in three steps. The first step is to estimate a predictive model for shorting fees that we can use to impute shorting fees over our entire sample. We use the 2010–2012 sample to estimate the following cross-sectional model for shorting fees:

$$Fee_{i,t} = b_{0,t} + b_{1,t}^{'} x_{i,t} + \epsilon_{i,t}$$

where x_{it} is a vector of bond characteristics. To check robustness, we present several models or specifications whereby we add in progressively more bond characteristics. Panel A of Table 12 runs monthly Fama-MacBeth regressions of bond shorting fee during 2010–2012 on bond issuance size decile dummies (Model 1), Model 1 augmented by on-the-run bond dummy (Model 2), Model 2 augmented by bond maturity dummies for $(0,1]$, $(1,2]$, $(2,3]$, $(3,5]$, $(5,7]$, $(7,10]$, 10+ years (Model 3), and Model 3 augmented by interactions between size, the on-the-run dummy, and maturity dummies (Model 4). Panel A reports the cross-sectional regression R^2. Not surprisingly, as we add in more bond characteristics, the R^2 rises.

[27] We have (in an omitted table available from the authors on request) also used the Wall Street Journal Survey and found similar results to the SPF. So it is unlikely that this finding is simply an artifact of the SPF survey. The literature finds that the three main surveys, SPF, Blue Chip, and Wall Street Journal, typically yield similar answers in empirical analyses (Schuh 2001).

[28] The table using unemployment disagreement is available from the authors on request.

Table 12
Disagreement about inflation, imputed shorting fee, and excess returns

Panel A. Fit of shorting fee imputation

	Model 1	Model 2	Model 3	Model 4
R^2	0.05	0.06	0.13	0.30

Panel B. Return-fee slope.
$r_{i,t} = c_1 \hat{Fee}_{i,t-1} + c_2 Maturity_{i,t-1} + c_0 + \epsilon_{i,t}$

	Model 1	Model 2	Model 3	Model 4
c_1	−29***	−27***	−21***	−8.1***
	(−4.3)	(−3.9)	(−6.7)	(−3.3)

Panel C. Return-fee slope on disagreement.
$\hat{c}_{1,t} = \nu_1 Disagreement_{t-1} + \nu_2 CP_{t-1} + \nu_3 LN_{t-1} + \nu_4 End\ of\ Recession_{t-1} + \nu_0 + \psi_t$

	Model 1	Model 2	Model 3	Model 4
ν_1 (no control)	−15***	−8.5***	−4.8***	−6.6***
	(−3.9)	(−2.9)	(−2.6)	(−3.5)
ν_1 (control *CP*)	−13***	−6.4***	−4.5**	−6.7***
	(−3.9)	(−3.1)	(−2.1)	(−3.5)
ν_1 (control *LN*)	−13***	−5.7*	−2.8	−6***
	(−3.1)	(−1.9)	(−1.6)	(−4)
ν_1 (control *End of Recession*)	−15***	−8.9***	−4.9***	−6.8***
	(−3.9)	(−3.1)	(−2.7)	(−3.7)
ν_1 (control all)	−13***	−5.6**	−3.1**	−6.5***
	(−3.5)	(−2.5)	(−2)	(−4.8)

Panel A runs monthly Fama-MacBeth regression of cross-sectional bond shorting fee during 2010–2012 on bond issuance size decile dummies (Model 1), Model 1 augmented by on-the-run dummy (Model 2), Model 2 augmented by bond maturity dummies for (0,1], (1,2], (2,3], (3,5], (5,7], (7,10], 10+ years (Model 3), and Model 3 augmented by interactions between size, on-the-run, and maturity dummies (Model 4). Panel A reports the cross-sectional regression R-squared. Imputed shorting fee *FeeImpute* is constructed for 1978–2012 using the regression coefficients in panel A. Panel B reports Fama-MacBeth regressions of 1-year holding excess bond returns on past-month imputed shorting fees. Panel C regresses the time-series slope coefficient of imputed shorting fee from panel B (coefficient c_1) on inflation disagreement during 1978–2012. Data on bond returns come from Gurkaynak, Sack, and Wright (2007) and are available on the FED Web site. Newey-West adjusted t-stats allowing for 13 lags are in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

We then use these regressions to predict the shorting fee that can be calculated over the whole sample and not just for the 2010–2012 period:

$$\hat{Fee}_{i,t} = \hat{b}_0 + \hat{b}_1' x_{i,t},$$

where $\hat{b}_0$ and $\hat{b}_1$ are the time series average of the coefficients from the monthly Fama-MacBeth regressions. Using this imputed or predicted fee $\hat{Fee}_{i,t}$, we then re-estimate Table 3:

$$r_{i,t} = c_{0,t} + c_{1,t} \hat{Fee}_{i,t-1} + c_{2,t} x_{i,t-1} + \epsilon_{i,t}.$$

Panel B of Table 12 reports the Fama-MacBeth estimates of the 1-year holding excess bond returns on past-month imputed shorting fees for specification (1) of Table 3, where the main bond characteristics control is $Maturity_{i,t-1}$. Similar to when we estimated Table 3 using actual fees over the 2010–2012 sample, we find a negative coefficient across all four specifications or models. The coefficients are all statistically significant. Moreover, the implied economic significance is also quite comparable across the four specifications: a one standard deviation increase in the predicted fee yields a decrease of expected returns that is around 6% of the standard deviation of the left hand side variable.

Finally, we in panel C regress the time-series slope coefficient on imputed shorting fee from panel B (coefficient $\hat{c}_{1,t}$) on inflation disagreement during 1978–2012:

$$\begin{aligned}\hat{c}_{1,t} =& \nu_1 Disagreement_{t-1} + \nu_2 CP_{t-1} + \nu_3 LN_{t-1} \\ &+ \nu_4 End\ of\ Recession_{t-1} + \nu_0 + \psi_t.\end{aligned}$$

We find that the coefficient ν_1 is negative—when inflation disagreement is higher, $\hat{c}_{1,t}$ is more negative, meaning that a portfolio that is long bonds that are costly to short, and short bonds that are cheap to short, earns lower returns when inflation disagreement is high. This is consistent with the importance of disagreement and shorting frictions in generating time-varying excess return predictability. This finding is robust across the different predictive models for shorting fee and various types of time series controls.

6. Conclusion

This paper reports consistent evidence that uncertainty about inflation, counter to the liquidity premium theory of interest rates, flattens the term structure of bond excess returns. We also find evidence to support that this puzzling finding is due to speculation in bond markets. When inflation uncertainty is high, investors disagree about expected inflation. As a result, optimistic investors expecting low inflation prefer to bet on long-term bonds, which are more sensitive to inflation than short-term ones. If short-sale restrictions keep pessimists investors sidelined, then disagreement leads to overpriced long-term maturities. We show that this inflation-betting channel interacts with the liquidity effect and flattens the term structure of bond returns, especially when the aggregate supply of bonds outstanding in the market is low.

Our inflation-betting effect has relevance for the current debate over the phenomenon of reach-for-yield. A number of prominent economists have argued that the low interest rates set by the Federal Reserve bank are potentially destabilizing since they encourage speculation by financial institutions in search of a minimum amount of yield. Yield is typically obtained by buying longer term bonds or duration in the context of Treasuries or to stretch for risk in other asset classes like junk bonds. Some initial evidence of reaching-for-yield by financial institutions is in Stein (2013) and Becker and Ivashina (2015), though these authors do not consider disagreement measures at all.

Our inflation-betting effect provides some support for such an hypothesis, to the extent that both mechanisms emphasize speculation in terms of bearing additional duration risk. If institutions reach for yield by speculating on long bonds, then their expectations about how long interest rates will stay at a low level becomes important. As a result, disagreement is also likely to play a role in explaining reach-for-yield behavior among institutions. Of course, in this context, disagreement about what the Federal Reserve Bank will do regarding its quantitative easing strategy is the key disagreement variable that should matter. Further work exploring the potential relationship between reaching-for-yield and disagreement about monetary policy more broadly would be a promising avenue for future research.

Appendix

A.1 Proof of Theorem 1

A.1.1 All investors hold a long portfolio of bonds. Consider the first case described in Theorem 1. In this equilibrium, investors from all three groups are long all maturity. Furthermore, assume that in this equilibrium, the yield of a bond of maturity k at date t is given by

$$\forall \tau \geq 0,\ \forall k \geq 1:\ \ k y_\tau^{(k)} = a_k \pi_\tau + b_k.$$

This will be later verified to hold in equilibrium.

With these assumptions, we can re-write the date-$t+1$ real wealth of investors in group $i \in \{A, B, a\}$ as[29]

$$\tilde{V}^{\,i}_{t+1} = W_t + \frac{1}{\Pi_t}\left(\sum_{k=2}^{K} x_{t,i}^{(k)}\left(k y_t^{(k)} - b_{k-1} - (1+a_{k-1})\tilde{\pi}_{t+1}\right) + x_{t,i}^{(1)}\left(y_t^{(1)} - \tilde{\pi}_{t+1}\right)\right). \tag{A1}$$

The approximation follows from a first-order Taylor expansion around zero inflation and yield.[30] Since investors have mean variance utility over their date-$t+1$ real wealth, the program of investors in group i is approximated at a first-order by:

$$\max_{x_{t,i}^{(k)}} \ \ \sum_{k=2}^{K} x_{t,i}^{(k)}\left(k y_t^{(k)} - (1+a_{k-1})\mathbb{E}_t^i[\tilde{\pi}_{t+1}] - b_{k-1}\right) + x_{t,i}^{(1)}\left(y_t^{(1)} - \mathbb{E}_t^i[\tilde{\pi}_{t+1}]\right)$$
$$- \frac{\sigma_\epsilon^2}{2\gamma}\left(x_{t,i}^{(1)} + \sum_{k=2}^{K}(1+a_{k-1})x_{t,i}^{(k)}\right)^2.$$

Since the equilibrium is assumed to have all investors long, the first-order conditions apply for all groups of investors: $\forall i \in \{a, A, B\}$,

$$\begin{cases} k y_t^{(k)} - \left(1+a_{k-1}\right)\left(\mu + \rho(\pi_t - \mu) + \lambda^i\right) - b_{k-1} = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(x_{t,i}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,i}^{(k)}(1+a_{k-1})\right) \\ y_t^{(1)} - (\mu + \rho(\pi_t - \mu) + \lambda^i) = \dfrac{\sigma_\epsilon^2}{\gamma}\left(x_{t,i}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,i}^{(k)}(1+a_{k-1})\right). \end{cases}$$

We can simply multiply the previous equations by the share of each group in the economy—$1-\alpha$ for HF, $\frac{\alpha}{2}$ for MF in group A and B—and then sum the first-order conditions across investors group, injecting the market

[29] Since we assume that HF are also long all bonds, we do not need to carry the shorting cost term c.

[30] This approximation neglects the second order terms, which would be of the order of the variance of the inflation rate and the yield. They are empirically small relative to the expected inflation rate and yield. For example, the average and the variance of the US monthly continuously compounded inflation rate is 0.0030 and 0.000012, respectively, between 1947 and 2014. The average and the variance of the US monthly 10-year Treasury constant maturity rate is 0.065 and 0.00077, respectively, between 1962 and 2014.

clearing condition of each bond maturity:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu+\rho(\pi_t-\mu)) - b_{k-1} = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1}) \left(Q_t^{(1)} + \displaystyle\sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1}) \right) \\ y_t^{(1)} - \mu - \rho(\pi_t-\mu) = \dfrac{\sigma_\epsilon^2}{\gamma} \left(Q_t^{(1)} + \displaystyle\sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1}) \right). \end{cases}$$

Remember that, by definition, for all t and k, the realized yield is given by $ky_t^{(k)} = a_k\pi_t + b_k$. By the previous equation, it is thus obvious that, for the equilibrium to be consistent, one needs: $a_1 = \rho$ and $a_k = \rho(1+a_{k-1})$ so that $a_k = \frac{\rho(1-\rho^k)}{1-\rho}$. Let $\Omega_t = \frac{\sigma_\epsilon^2}{\gamma}\left(\sum_{k=1}^{K} \frac{1-\rho^k}{1-\rho} Q_t^{(k)}\right)$ be the aggregate weighted supply of bonds. By using the value of a_k, we can recover the constant in the yield of the short bond: $b_1 = (1-\rho)\mu + \frac{\sigma_\epsilon^2}{\gamma} S$. Note from the first equation in the previous system of equations that: $b_k = b_{k-1} + \frac{1-\rho^k}{1-\rho} b_1$. This trivially implies that:

$$b_k = \left(\sum_{l=1}^{k} \frac{1-\rho^l}{1-\rho} \right) b_1.$$

Now that we have found the price of all outstanding bonds, we need to check whether these prices correspond indeed to an equilibrium, that is, all agents are in fact long all bonds in this equilibrium. Since there is only one source of risk in the economy, the holdings of agents are not pinned down in equilibrium, only their weighted exposure to inflation risk (i.e., $x_{t,i}^{(1)} + \sum_{k=2}^{K} x_{t,i}^{(k)}(1+a_{k-1})$ in the previous equation). From the first-order condition on the short rate, we see that the pessimists have a long portfolio of bonds if and only if:

$$y_t^{(1)} - (\mu+\rho(\pi_t-\mu)+\lambda) > 0 \Leftrightarrow \lambda < \frac{\sigma_\epsilon^2}{\gamma}\Omega_t.$$

Provided disagreement is not too large, even pessimist investors remain long bonds and share some inflation risk with the pessimists. In this case, expected nominal log returns are given by

$$\begin{aligned} \mathbb{E}\left[\tilde{R}_t^{(k)}\right] &= \mathbb{E}\left[\log\left(\tilde{P}_{t+1}^{(k-1)}\right)\right] - \log\left(P_t^{(k)}\right) = ky_t^{(k)} - \mathbb{E}_t\left[(k-1)\tilde{y}_{t+1}^{(k-1)}\right] \\ &= \mu + \rho(\pi_t-\mu) + \left(\frac{1-\rho^k}{1-\rho}\right)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t. \end{aligned}$$

And yields are given by:

$$y_t^{(k)} = \frac{a_k}{k}\pi_t + \frac{b_k}{k} = \mu + \left(\frac{1-\rho^k}{k(1-\rho)}\right)\rho(\pi_t-\mu) + \frac{1}{k}\left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t$$

A.1.2 HFs hold a long portfolio of all bonds; pessimist MFs are sidelined. Consider now the second equilibrium described in Theorem 1. In this equilibrium, pessimists MFs are sidelined from the bond market, but HFs, nonetheless, hold a long portfolio of bonds of all maturities. Assume again that in this equilibrium:

$$\forall \tau \geq 0, \forall k \geq 1: \; ky_\tau^{(k)} = a_k\pi_\tau + b_k$$

In this equilibrium, the optimists' first-order conditions will hold:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu+\rho(\pi_t-\mu)-\lambda) - b_{k-1} = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1}) \left(x_{t,A}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,A}^{(k)}(1+a_{k-1}) \right) \\ y_t^{(1)} - (\mu+\rho(\pi_t-\mu)-\lambda) = \dfrac{\sigma_\epsilon^2}{\gamma} \left(x_{t,A}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,A}^{(k)}(1+a_{k-1}) \right), \end{cases}$$

while the HFs' first-order conditions become:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu+\rho(\pi_t-\mu)) - b_{k-1} = \frac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(x_{t,a}^{(1)} + \sum_{k=2}^{K} x_{t,a}^{(k)}(1+a_{k-1})\right) \\ y_t^{(1)} - (\mu+\rho(\pi_t-\mu)) = \frac{\sigma_\epsilon^2}{\gamma}\left(x_{t,a}^{(1)} + \sum_{k=2}^{K} x_{t,a}^{(k)}(1+a_{k-1})\right). \end{cases}$$

We can multiply the first system by $\frac{\alpha}{2}$, multiply the second system by $1-\alpha$, sum them, and injecting the market-clearing condition that $\frac{\alpha x_{t,A}^{(k)}}{2} + (1-\alpha)x_{t,a}^{(k)} = Q_t^{(k)}$.

$$\begin{cases} (1-\frac{\alpha}{2})\left(ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu+\rho(\pi_t-\mu)) - b_{k-1}\right) + \frac{\alpha}{2}(1+a_{k-1})\lambda \\ = \frac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(Q_t^{(1)} + \sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1})\right) \\ (1-\frac{\alpha}{2})\left(y_t^{(1)} - (\mu+\rho(\pi_t-\mu))\right) + \frac{\alpha}{2}\lambda = \frac{\sigma_\epsilon^2}{\gamma}\left(Q_t^{(1)} + \sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1})\right) \end{cases}$$

We can rewrite this system as:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu+\rho(\pi_t-\mu)) - b_{k-1} + \theta\left(1+a_{k-1}\right)\lambda = (1+\theta)\frac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\Omega_t \\ y_t^{(1)} - \mu - \rho(\pi_t-\mu) + \theta\lambda = (1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t. \end{cases}$$

We see again that $a_1 = \rho$ and for all $k \geq 2$, $a_k = (1+a_{k-1})\rho$. This implies that for all $k \geq 1$, $a_k = \frac{\rho(1-\rho^k)}{1-\rho}$. We also derive from the second equation in the previous system that:

$$b_1 = (1-\rho)\mu + (1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda.$$

The first equation of the previous system then implies $b_k = b_{k-1} + \frac{1-\rho^k}{1-\rho}b_1$, which trivially implies that

$$b_k = \left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)b_1 = \left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)\left((1-\rho)\mu + (1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right)$$

For this to be an equilibrium, it needs to be that arbitrageurs are, in fact, long in equilibrium. This condition is simply that:

$$b_1 > (1-\rho)\mu \Leftrightarrow \frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t > \lambda.$$

It also needs to be the case that pessimists would, in fact, want to short all bonds if they could, which is equivalent to their marginal utility at zero-holding being negative:

$$\lambda > b_1 - \mu(1-\rho) \Leftrightarrow \lambda > \frac{\sigma_\epsilon^2}{\gamma}\Omega_t.$$

Expected returns in this equilibrium are given by:

$$\mathbb{E}[\tilde{R}_t^{(k)}] = ky_t^{(k)} - \mathbb{E}_t\left[(k-1)\tilde{y}_{t+1}^{(k-1)}\right] = \rho\pi_t + (1-\rho)\mu + \left(\frac{1-\rho^k}{1-\rho}\right)\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right).$$

And yields are given by:

$$y_t^{(k)} = \frac{a_k}{k}\pi_t + \frac{b_k}{k} = \mu + \left(\frac{1-\rho^k}{k(1-\rho)}\right)\rho(\pi_t-\mu) + \frac{1}{k}\left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right).$$

A.1.3 HFs short the bond markets; pessimist MFs are sidelined. We now consider the third equilibrium described in Theorem 1. In this equilibrium, pessimists are sidelined from the bond market and HFs are shorting bonds. Assume again that in this equilibrium:

$$\forall \tau \geq 0, \forall k \geq 1: \; ky_\tau^{(k)} = a_k \pi_\tau + b_k.$$

In this equilibrium, the optimists' first-order conditions will hold:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu + \rho(\pi_t - \mu) - \lambda) - b_{k-1} = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(x_{t,A}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,A}^{(k)}(1+a_{k-1})\right) \\ y_t^{(1)} - (\mu + \rho(\pi_t - \mu) - \lambda) = \dfrac{\sigma_\epsilon^2}{\gamma}\left(x_{t,A}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,A}^{(k)}(1+a_{k-1})\right). \end{cases}$$

The HFs' first-order condition has to be modified to account for HFs short bonds of all maturities:

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu + \rho(\pi_t - \mu)) - b_{k-1} = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(x_{t,a}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,a}^{(k)}(1+a_{k-1})\right) + cx_{t,a}^{(k)} \\ y_t^{(1)} - (\mu + \rho(\pi_t - \mu)) = \dfrac{\sigma_\epsilon^2}{\gamma}\left(x_{t,a}^{(1)} + \displaystyle\sum_{k=2}^{K} x_{t,a}^{(k)}(1+a_{k-1})\right) + cx_{t,a}^{(1)} \end{cases}$$

We can multiply the first system by $\frac{\alpha}{2}$, multiply the second system by $1-\alpha$, and sum them, injecting the market-clearing condition that $\frac{\alpha x_{t,A}^{(k)}}{2} + (1-\alpha)x_{t,a}^{(k)} = Q_t^{(k)}$. Introduce $\theta = \frac{\frac{\alpha}{2}}{1-\frac{\alpha}{2}}$:

$$\begin{cases} (1-\dfrac{\alpha}{2})\left(ky_t^{(k)} - (1+a_{k-1})(\mu + \rho(\pi_t - \mu)) - b_{k-1}\right) + \dfrac{\alpha}{2}(1+a_{k-1})\lambda \\ = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\left(Q_t^{(1)} + \displaystyle\sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1})\right) + (1-\alpha)cx_{t,a}^{(k)} \\ (1-\dfrac{\alpha}{2})\left(y_t^{(1)} - (\mu + \rho(\pi_t - \mu))\right) + \dfrac{\alpha}{2}\lambda = \dfrac{\sigma_\epsilon^2}{\gamma}\left(Q_t^{(1)} + \displaystyle\sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1})\right) + (1-\alpha)cx_{t,a}^{(1)} \end{cases}$$

Define $\Omega_t = Q_t^{(1)} + \sum_{k=2}^{K} Q_t^{(k)}(1+a_{k-1})$. We can rewrite this system as

$$\begin{cases} ky_t^{(k)} - \left(1+a_{k-1}\right)(\mu + \rho(\pi_t - \mu)) - b_{k-1} + \theta\left(1+a_{k-1}\right)\lambda = (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\Omega_t + c(1-\theta)x_{t,a}^{(k)} \\ y_t^{(1)} - \mu - \rho(\pi_t - \mu) + \theta\lambda = (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + (1-\theta)cx_{t,a}^{(1)}. \end{cases}$$

We can plug these equilibrium conditions into HFs first-order conditions:

$$\begin{cases} (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\Omega_t - \theta\left(1+a_{k-1}\right)\lambda = \dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})S_t^a + c\theta x_{t,a}^{(k)} \\ (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda = \dfrac{\sigma_\epsilon^2}{\gamma}S_t^a + c\theta x_{t,a}^{(1)}, \end{cases}$$

where $S_t^a = \left(x_{t,a}^{(1)} + \sum_{k=2}^{K} x_{t,a}^{(k)}(1+a_{k-1})\right)$. To eliminate S_t^a from the previous equations, we can simply multiply the equations for $k > 1$ by $(1+a_{k-1})$ and sum all these equations. Define $\Theta = 1 + \sum_{k=2}^{K}(1+a_{k-1})^2$, and we then have:

$$S_t^a = \Theta \frac{(1+\theta)\Omega_t \frac{\sigma_\epsilon^2}{\gamma} - \lambda\theta}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}.$$

We can then derive the expression for $x_{t,a}^{(k)}$ from the previous system:

$$\begin{cases} x_{t,a}^{(k)} = (1+a_{k-1})\dfrac{\frac{\sigma_\epsilon^2}{\gamma}(1+\theta)\Omega_t - \lambda\theta}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta} \\ x_{t,a}^{(1)} = \dfrac{\frac{\sigma_\epsilon^2}{\gamma}(1+\theta)\Omega_t - \lambda\theta}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}. \end{cases}$$

We can then use these expressions to substitute in the system defining the equilibrium yields and obtain:

$$\begin{cases} ky_t^{(k)} - (1+a_{k-1})(\mu + \rho(\pi_t - \mu)) - b_{k-1} + \theta(1+a_{k-1})\lambda \\ = (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}(1+a_{k-1})\Omega_t + c(1-\theta)(1+a_{k-1})\dfrac{\frac{\sigma_\epsilon^2}{\gamma}(1+\theta)\Omega_t - \lambda\theta}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta} \\ y_t^{(1)} - \mu - \rho(\pi_t - \mu) + \theta\lambda = (1+\theta)\dfrac{\sigma_\epsilon^2}{\gamma}\Omega_t + (1-\theta)c\dfrac{\frac{\sigma_\epsilon^2}{\gamma}(1+\theta)\Omega_t - \lambda\theta}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}. \end{cases}$$

We see again that $a_1 = \rho$ and for all $k \geq 2$, $a_k = (1+a_{k-1})\rho$. This implies that for all $k \geq 1$, $a_k = \frac{\rho(1-\rho^k)}{1-\rho}$. We also get: $\Theta = \sum_{k=1}^{K}\left(\frac{1-\rho^k}{1-\rho}\right)^2$ and $\Omega_t = \left(\sum_{l=1}^{K}\frac{1-\rho^l}{1-\rho}Q_t^{(l)}\right)$. We have

$$b_1 = \mu(1-\rho) + \frac{\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega - \theta\lambda\right)\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}.$$

We also have that: $b_k = b_{k-1} + \frac{1-\rho^k}{1-\rho}b_1$, which trivially implies that:

$$b_k = \left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)b_1 = \left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)\left(\mu(1-\rho) + \frac{\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right)\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}\right).$$

For this to be an equilibrium, it needs to be that arbitrageurs are in fact short in equilibrium. This condition is simply that:

$$\lambda > \frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t.$$

It also needs to be the case that pessimists would in fact want to short all bonds if they could, which is equivalent to:

$$\lambda > b_1 - \mu(1-\rho) \Leftrightarrow \lambda > \frac{\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right)\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta} \Leftrightarrow \lambda > \frac{\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)}{(1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Theta + 2c\theta}\frac{\sigma_\epsilon^2}{\gamma}(1+\theta)\Omega_t.$$

Note that if $\lambda > \frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t$, then the previous condition is immediately verified since $(1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Theta + 2c\theta > \theta\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)$. Expected returns in this equilibrium are given by

$$\mathbb{E}[\tilde{R}_t^{(k)}] = ky_t^{(k)} - \mathbb{E}_t\left[(k-1)\tilde{y}_{t+1}^{(k-1)}\right] = \mu + \rho(\pi_t - \mu) + \left(\frac{1-\rho^k}{1-\rho}\right)\frac{\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right)\left(\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\right)}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}.$$

Finally, yields are simply given by

$$
\begin{aligned}
y_t^{(k)} &= \frac{a_k}{k}\pi_t + \frac{b_k}{k} \\
&= \mu + \left(\frac{1-\rho^k}{k(1-\rho)}\right)\rho(\pi_t - \mu) \\
&\quad + \frac{1}{k}\left(\sum_{l=1}^{k}\frac{1-\rho^l}{1-\rho}\right)\left(\frac{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}\right)\left((1+\theta)\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \theta\lambda\right).
\end{aligned}
$$

A.2 Proof of Corollary 1

From Theorem 1, we know that:

$$
\begin{aligned}
\mathbb{E}[\tilde{R}_t^{(k)}] &= \mu + \rho(\pi_t - \mu) \\
&\quad + \left(\frac{1-\rho^k}{1-\rho}\right)\left(\frac{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c}{\frac{\sigma_\epsilon^2}{\gamma}\Theta + c\theta}\right)^{1\left\{\lambda > \frac{\sigma_\epsilon^2}{\gamma}\frac{1+\theta}{\theta}\Omega_t\right\}}\left(\frac{\sigma_\epsilon^2}{\gamma}\Omega_t - \mathbb{1}_{\left\{\lambda > \frac{\sigma_\epsilon^2}{\gamma}\Omega_t\right\}}\theta\left(\lambda - \frac{\sigma_\epsilon^2}{\gamma}\Omega_t\right)\right).
\end{aligned}
$$

The results in Corollary 1 directly follow from this expression of bonds expected returns.

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