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Citation

Hanson, Samuel G., David O. Lucca, and Jonathan H. Wright. "Rate-Amplifying Demand and the Excess Sensitivity of Long-Term Rates." *Quarterly Journal of Economics* 136, no. 3 (August 2021): 1719–1781.

Published Version

<https://doi.org/10.1093/qje/qjab011>

Permanent link

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RATE-AMPLIFYING DEMAND AND THE EXCESS SENSITIVITY OF LONG-TERM RATES*

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Abstract

Long-term nominal interest rates are surprisingly sensitive to high-frequency (daily or monthly) movements in short-term rates. Since 2000, this high-frequency sensitivity has grown even stronger in U.S. data. By contrast, the association between low-frequency changes (at 6- or 12-month horizons) in long- and short-term rates, which was also strong before 2000, has weakened substantially. This puzzling post-2000 pattern arises because increases in short rates temporarily raise the term premium component of long-term yields, leading long rates to *temporarily* overreact to changes in short rates. The frequency-dependent excess sensitivity of long-term rates that we observe in recent years is best understood using a model in which (i) declines in short rates trigger “rate-amplifying” shifts in investor demand for long-term bonds and (ii) the arbitrage response to these demand shifts is both limited and slow. We study, both theoretically and empirically, how such rate-amplifying demand can be traced to mortgage re-financing activity, investors who extrapolate recent changes in short rates, and investors who “reach for yield” when short rates fall. We discuss the implications of our findings for the validity of event-study methodologies and the transmission of monetary policy.

JEL codes: E43, E52, G12, G14.

*Earlier versions of this paper circulated under the title “The Excess Sensitivity of Long-Term Rates: A Tale of Two Frequencies” and “Interest Rate Conundrums in the Twenty First Century” We thank John Campbell, Anna Cieslak, Gabriel Chodorow-Reich, Richard Crump, Thomas Eisenbach, Robin Greenwood, Andrei Shleifer, Eric Swanson, Jeremy Stein, and Adi Sunderam as well as seminar participants at the 2018 AEA meetings, Chicago Booth, Harvard University, the Federal Reserve Bank of New York, Johns Hopkins University, Northwestern Kellogg, the Society for Computational Economics 2017 International Conference, University of Georgia, and University of Southern California and four anonymous referees for helpful comments. Hanson gratefully acknowledges funding from the Division of Research at Harvard Business School. All errors are our sole responsibility. The views expressed here are the authors’ and are not representative of the views of the Federal Reserve Bank of New York or of the Federal Reserve System. Emails: shanson@hbs.edu; david.lucca@ny.frb.org; wrightj@jhu.edu.

I. INTRODUCTION

The sensitivity of long-term interest rates to movements in short-term rates is a central feature of the term structure and plays a crucial role in the transmission of monetary policy. Short-term nominal interest rates are determined by current monetary policy and its near-term expected path. Shocks to real interest rates are generally thought to be short-lived, so long-term nominal rates should not be highly sensitive to changes in nominal short rates if long-run inflation expectations are well anchored and the expectations hypothesis holds (Shiller, 1979). In fact, long-term nominal rates are surprisingly sensitive to high-frequency changes in short rates (Shiller et al., 1983; Cochrane and Piazzesi, 2002; Gürkaynak et al., 2005; Giglio and Kelly, 2018). The deeper forces underpinning this puzzling degree of high-frequency sensitivity, and its evolution, remain poorly understood.

We provide evidence that, in the past two decades, the sensitivity of long rates has grown even stronger at high frequencies (daily or monthly), but has weakened substantially at lower frequencies (6 or 12 month horizons)¹. As a result, the sensitivity of long rates to changes in short rates has become highly frequency-dependent since 2000. Between 1971 and 1999, a daily regression of changes in 10-year U.S. Treasury yields on changes in 1-year yields delivers a coefficient of 0.56; and the analogous regression using 12-month changes gives nearly the same coefficient. Between 2000 and 2019, the coefficient from the daily regression jumps to 0.87, while the coefficient from the corresponding 12-month regression drops to 0.23. Figure I, which plots the sensitivity of 10-year yields to changes in 1-year yields as a function of horizon in both the pre-2000 and post-2000 samples, summarizes this key finding. This pattern is not specific to the U.S.: we find similar results for Canada, Germany, and the U.K.

What explains this puzzling post-2000 tendency of short- and long-term rates to move together at high but not low frequencies? Statistically, this pattern arises because recent increases in short rates predict a subsequent flattening of the yield curve, as well as declines in long-term yields and forward rates, in the post-2000 data. These predictable reversals in long rates are linked to a new form of short-lived bond return predictability: since 2000, the expected returns on long-term bonds (in excess of those on short-term bonds) are *temporarily* elevated following increases in short rates. Thus, relative to an expectations-hypothesis baseline, long rates temporarily overreact to changes in short rates, exhibiting what Mankiw and Summers (1984) dubbed “excess sensitivity.” In the post-2000 data, we estimate that 10-year yields rise by 66 basis points in response to a 100 bps monthly increase in 1-year yields. Over the next 6 months, 10-year yields are expected to fall by 36 bps, reversing over half of the initial response.

What deeper forces underpin the evolving sensitivity of long rates to movements in short rates?

¹We do not mean to argue that there was a *discrete* change in the underlying data-generating process around 2000. Instead, our reading of the evidence is that the underlying data-generating process has changed *gradually* over time.

Gürkaynak et al. (2005) note that the strong sensitivity of long-term nominal rates could be consistent with the expectations hypothesis if one adopts the view that long-run inflation expectations are unanchored and are being continuously updated—i.e., if there are large shocks to trend inflation as in Stock and Watson (2007). This narrative is a good explanation for the high degree of sensitivity observed before 2000. Consistent with the expectations-hypothesis logic of this explanation, in the pre-2000 data, we find no evidence that the reaction of long yields to movements in short rates tends to predictably reverse.

However, in the post-2000 period, the high-frequency sensitivity of long-term nominal rates primarily reflects the sensitivity of long-term *real* rates to nominal short rates, rather than the sensitivity of break-even inflation (Beechey and Wright, 2009; Hanson and Stein, 2015). To the extent that one shares the widespread view that expected future real rates at distant horizons do not fluctuate meaningfully at high frequencies (see Gürkaynak et al., 2005), this makes it hard to square the strong post-2000 sensitivity at high frequencies with expectations-hypothesis logic. Hanson and Stein (2015) argue that the strong post-2000 sensitivity works through the term premium component of long-term yields: shocks to short rates move term premia in the same direction. Consistent with this view, we find that the reaction of long yields to movements in short rates tends to predictably reverse in the post-2000 data, giving rise to short-lived shifts in the expected returns to holding long-term bonds.

How can we best understand our finding that, in recent decades, the sensitivity of long rates to changes in short rates declines steeply with horizon? Because it reflects a form of short-lived return predictability, the most natural explanations involve temporary supply-and-demand imbalances in financial markets (De Long et al. (1990), Shleifer and Vishny (1997), and Duffie (2010)). We develop a model of these imbalances that emphasizes the role of what we call “rate-amplifying” shocks to the supply-and-demand for long-term bonds. Our model builds on Greenwood and Vayanos (2014) and Vayanos and Vila (2020) who stress the limited risk-bearing capacity of the specialized fixed-income arbitrageurs who must absorb shocks to the net supply of long-term bonds. In our model, risk-averse bond arbitrageurs can invest in either short- or long-term nominal bonds. While monetary policy pins down the interest rate on short-term bonds, long-term bonds are available in a net supply that varies over time. This net supply, which arbitrageurs must hold in equilibrium, equals the gross supply of long-term bonds net of the amount inelastically demanded by other, non-arbitrageur investors. To induce risk-averse arbitrageurs to absorb an increase in net supply of long-term bonds, the expected return on long-term bonds in excess of that on short-term bonds must rise, lifting the term premium component of long-term yields.

Our explanation adds two novel ingredients to this familiar setup: (i) “rate-amplifying” shifts in the supply or demand for long-term bonds and (ii) a slow-moving arbitrage response. First, we assume that shocks to the net supply of long-term bonds are *positively* correlated with shocks to

short rates. This positive correlation can obtain either because increases in short rates are associated with increases in the gross supply of long-term bonds or with non-standard reductions in the demand of other, non-arbitrageur investors. Since arbitrageurs' risk-bearing capacity is limited, this implies that increases in short rates are associated with increases in the term premium component of long rates, generating "excess sensitivity" relative to the expectations hypothesis. This reduced-form assumption is consistent with several distinct rate-amplification mechanisms that we detail below, each rooted in well-known institutional frictions and facets of investor psychology, that have arguably grown in importance recent decades.

The second ingredient is that arbitrage capital is slow-moving as in [Duffie \(2010\)](#). As a result, these rate-amplifying demand shocks encounter a short-run arbitrage demand curve that is steeper than the long-run arbitrage demand curve, generating a short-lived imbalance in the market for long-term bonds. This slow-moving capital dynamic implies that the shifts in term premia triggered by movements in short rates are transitory. As a result, the excess sensitivity of long rates is greatest when measured at high frequencies. Furthermore, we show that frequency-dependent excess sensitivity is most pronounced when the underlying rate-amplifying demand shocks are themselves short-lived. In summary, the combination of rate-amplifying demand shocks and slow-moving arbitrage capital enables our model to match the frequency-dependent sensitivity of long rates observed since 2000.

We explore three rate-amplification channels: (i) shifts in the effective gross supply of long-term bonds due to mortgage refinancing waves ([Hanson, 2014](#); [Malkhozov et al., 2016](#)), (ii) shifts in the demand for long-term bonds from biased investors who extrapolate recent changes in short rates ([Giglio and Kelly, 2018](#); [D'Arienzo, 2020](#)), and (iii) shifts in the demand from investors who "reach for yield" when short rates fall ([Hanson and Stein, 2015](#)). For each channel, we first show how it can be used to microfound rate-amplifying shocks to the net supply of long-term bonds similar to those we previously assumed in reduced-form. Next, we discuss why the strength of each channel may have grown in recent decades: the key underlying trend here is the increasing financialization of interest-rate risk. Finally, by looking at the relationship between bond yields and different financial quantities, we empirically assess the extent to which each channel contributes to the frequency-dependent sensitivity of long-term rates we observe since 2000. We find evidence that mortgage refinancing and investor extrapolation both help explain why long yields rates have temporarily overreacted to short rates since 2000 in the U.S. By contrast, we find less evidence that reaching-for-yield plays an important role in driving our key empirical findings.

A vast literature demonstrates that, contrary to the expectations hypothesis, the expected excess returns on long-term bonds vary meaningfully over time ([Fama and Bliss, 1987](#); [Campbell and Shiller, 1991](#); [Cochrane and Piazzesi, 2005](#)). In contrast to most of the existing literature on bond return predictability which focuses on business-cycle frequency variation in expected returns, our

findings point to a new, short-lived form of return predictability that has emerged in recent decades.

Our findings have important implications for how economists should interpret event-study evidence based on high-frequency changes in long-term bond yields. Macroeconomic news, including news about monetary policy, is lumpy, and the short-run change in long-term rates around news announcements is often used as a measure of the expected longer-run impact of news shocks on short rates. [Nakamura and Steinsson \(2018\)](#) is a prominent recent example of this increasingly popular approach to identification in macroeconomics. But, if some of the impact of a news shock on long-term rates wears off quickly over time, then a shock’s short- and long-run impact are quite different. And event-studies only captures the short-run impact. For instance, it is common for news announcements to cause large jumps in 10-year forward rates, but since a large portion of these jumps are due to *transient* shifts in term premia, event-studies are likely to provide biased estimates of the longer-run impact on short rates.

Our results also have implications for monetary policy transmission. In the textbook New Keynesian view ([Gali, 2008](#)), the central bank adjusts short-term nominal rates. This affects long-term rates via the expectations hypothesis, which in turn influences aggregate demand. [Stein \(2013\)](#) points out that the excess sensitivity of long-term yields, whereby shocks to short rates move term premia in the same direction, should strengthen the effects of monetary policy relative to the textbook view. [Stein \(2013\)](#) refers to this as the “recruitment” channel of monetary transmission. We find that the behavior of interest rates does not conform to the textbook New Keynesian view in which term premia are constant. Nonetheless, our findings suggest that the recruitment channel may not be as strong as [Stein \(2013\)](#) speculates since a portion of the resulting shifts in term premia are transitory and, thus, likely to have only modest effects on aggregate demand. We do not argue that there is no recruitment channel, just that it is smaller than one might conclude based on the high-frequency response of term premia to policy shocks documented in [Hanson and Stein \(2015\)](#), [Gertler and Karadi \(2015\)](#), and [Gilchrist et al. \(2015\)](#).

In Section II., we document our key stylized facts about the changing high- and low-frequency sensitivity of long-term interest rates. In Section III., we show that past increases in short rates predict a future reversals in long-term yields in the post-2000 data, reflecting a new form of bond return predictability. Section IV. develops the model we use to interpret our findings. We build on this framework in Section V. where we explicitly microfound three specific rate-amplification mechanisms—mortgage refinancing, extrapolation, and reaching-for-yield—and then assess empirically the extent to which each mechanism helps explain our key findings. Section VI. discusses the implications of our findings for event-study identification strategies that exploit high-frequency movements in long-term yields, the transmission of monetary policy, bond market “conundrums,” and affine term structure models. Section VII. concludes.

II. THE SENSITIVITY OF LONG-TERM RATES TO SHORT-TERM RATES

Between 1971 and 2000, the sensitivity of long-term rates to changes in short-term rates was similarly strong at both high- and low-frequencies. Since 2000, the association between high-frequency changes in short- and long-term interest rates has grown even stronger. By contrast, the association between low-frequency changes in short- and long-term rates has weakened substantially. As a result, the sensitivity of long-term rates has become surprisingly frequency-dependent since 2000. We first document these basic facts for the U.S. We then contrast the patterns we see in the post-2000 data with those observed in the U.S. prior to the 1970s. Finally, we show that the sensitivity of long-term rates has evolved in a similar fashion in Canada, Germany, and the U.K.

Baseline findings for the U.S. We begin by regressing changes in 10-year Treasury yields or forward rates on changes in 1-year nominal Treasury yields. Specifically, we estimate regressions of the form:

$$(1) \quad y_{t+h}^{(10)} - y_t^{(10)} = \alpha_h + \beta_h(y_{t+h}^{(1)} - y_t^{(1)}) + \varepsilon_{t,t+h}$$

and

$$(2) \quad f_{t+h}^{(10)} - f_t^{(10)} = \alpha_h + \beta_h(y_{t+h}^{(1)} - y_t^{(1)}) + \varepsilon_{t,t+h},$$

where $y_t^{(n)}$ is the continuously compounded n -year zero-coupon yield in period t and $f_t^{(n)}$ is the n -year-ahead instantaneous forward rate. We obtain data on the nominal and real U.S. Treasury yield curve from [Gürkaynak et al. \(2007\)](#) and [Gürkaynak et al. \(2010\)](#). We decompose nominal yields into real yields and inflation compensation, defined as the difference between nominal and real yields derived from Treasury Inflation-Protected Securities (TIPS). Our sample begins in August 1971, which is when reliable data on 10-year nominal yields first become available and ends in December 2019. For real yields and inflation compensation, we only study the post-2000 sample, since data on TIPS are not available until 1999. All data are measured as of the end of the relevant period—e.g., the last trading day of each month.

In standard monetary economics models, the central bank sets overnight nominal interest rates, and other rates are influenced by the expected path of overnight rates. A large literature argues that central banks in the U.S. and abroad have increasingly relied on communication—implicit or explicit signaling about the future path of overnight rates—as an active policy instrument ([Gürkaynak et al., 2005](#)). To capture news about the near-term path of policy that would not impact the current overnight rate, we take the short rate to be the 1-year nominal Treasury rate which follows approaches in the recent literature ([Gertler and Karadi, 2015](#); [Gilchrist et al., 2015](#); [Hanson and Stein, 2015](#)).

Panel A in Table I reports estimated coefficients β_h in equation (1) for zero-coupon nominal yields, real yields, and inflation compensation using daily data and end-of-month data with $h = 1, 3, 6, 12$ months—i.e., for daily, monthly, quarterly, semi-annual, and annual changes in yields.² The results are shown for the pre-2000 and post-2000 samples separately. We base this sample split on a number of break-date tests that we will discuss shortly. Figure I plots the estimated coefficients β_h in equation (1) for nominal yields versus monthly horizon h for the pre-2000 and post-2000 samples.

Since we use overlapping h -month changes in equation (1) when $h > 1$, we report Newey and West (1987) standard errors using a lag truncation parameter of $\lceil 1.5 \times h \rceil$; when $h = 1$, we report heteroskedasticity-robust standard errors. To address the tendency for statistical tests based on Newey and West (1987) standard errors to over-reject in finite samples, we compute p -values using the asymptotic theory of Kiefer and Vogelsang (2005) which gives more conservative p -values and has better finite-sample properties than traditional Gaussian asymptotic theory.

Panel A shows that, prior to 2000, there was a strong tendency for short- and long-term yields to rise and fall together at both high- and low-frequencies. While the high-frequency relationship has grown even stronger since 2000, the low-frequency relationship has weakened significantly. Specifically, the daily coefficient for 10-year yields has risen from $\beta_{day} = 0.56$ in the pre-2000 sample to $\beta_{day} = 0.87$ in the post-2000 sample and this increase is statistically significant (p -val < 0.001). By contrast, the coefficient for $h = 12$ -month changes in 10-year yields has fallen from $\beta_{12} = 0.56$ before 2000 to $\beta_{12} = 0.23$ in the post-2000 sample and this decline is also highly significant (p -val < 0.001).

Combining these observations, Figure I shows our main finding: in the post-2000 sample, the β_h coefficients decline steeply with the horizon h . By contrast, β_h is a relatively constant function of h in the pre-2000 sample. Furthermore, Table I shows that the majority of the decline in β_h as a function of h in the post-2000 sample is due to the real component of long-term yields.

This is a surprising result: one would not expect β_h to vary strongly with monthly horizon h as in the post-2000 data. In a standard term-structure models with a single factor, we have $y_t^{(10)} = \alpha + \beta \cdot y_t^{(1)}$ for some $\beta \in (0, 1)$, implying that $\beta_h = \beta$ for all h , regardless of whether the expectations hypothesis holds. More generally, even accounting for multiple risk factors, term premia only fluctuate at business-cycle frequencies in conventional asset-pricing models, implying that β_h should be quite stable across monthly horizons h . And, as detailed in Section IV. below, if there are both persistent and transient shocks to short rates, the expectations hypothesis implies that β_h should be slightly increasing in h as it was in the pre-2000 data. Thus, our finding that β_h is a steeply decreasing function of horizon h since 2000 suggests that term structure dynamics

²Bond maturities are in years and time periods are in months, except when we estimate regressions at a daily frequency.

have shifted in an important way.

Panel B of Table I reports the corresponding β_h coefficients in equation (2) using changes in instantaneous forwards as the dependent variable. Like 10-year yields, the sensitivity of 10-year forward rates to changes in short-term rates has risen at high frequencies, but has declined markedly at low frequencies.

We use two approaches to date the timing of the break and both approaches suggest that there was a break around 2000. First, we estimate equations (1) and (2) using 10-year rolling windows. The estimated coefficients for $h = 12$ -month changes are shown in Figure II for 10-year yields and forwards. These β_{12} coefficients decline substantially in more recent windows. The second approach is to test for a structural break in equations (1) and (2) for $h = 12$ -month changes, allowing for an unknown break date. We use the test of Andrews (1993) who conducts a Chow (1960) test at all possible break dates, and then takes the maximum of the Wald test statistics. Figure III plots the Wald test statistic for each possible break date in equations (1) and (2) along with the Cho and Vogelsang (2017) critical values for a null of no structural break. The strongest evidence for a break is in 1999 or 2000 in both equations (1) and (2) and the break is highly statistically significant.³

To clarify, we do not intend to argue that there was a *discrete* change in the underlying data-generating process in 2000. Instead, consistent with the rolling-window regressions shown in Figure II, our reading of the data is that the underlying data-generating process has changed gradually over time—a gradual change which then becomes discernible when we compare the behavior of yields in across different samples. Nonetheless, throughout the remainder of the paper, we will adopt the heuristic of simply splitting the data into two samples: pre- and post-2000.

Robustness. In the Internet Appendix, we conduct a battery of robustness checks on our key findings. First, we show that similar results obtain if we use long-term private yields as the dependent variable in equation (1). We examine long-term corporate bond yields with Moody's ratings of Aaa and Baa, the 10-year swap yield, and the yield on mortgage-backed-securities. For all of these long rates, the sensitivity to changes in 1-year Treasury rates was similar irrespective of frequency before 2000. After 2000, the sensitivity at high frequencies increases while the sensitivity at low frequencies declines significantly.

Second, we obtain similar results using different proxies for the short-term rate—i.e., using changes in 3-month, 6-month, or 2-year Treasury yields—as the independent variable in equation (1).

Third, one might be concerned about our use of overlapping changes in equations (1) and (2)

³In the Internet Appendix, we date the emergence of the frequency-dependent sensitivity of long rates. We report 10-year rolling estimates of $\beta_1 - \beta_{12}$ for 10-year yields and find that $\beta_1 - \beta_{12}$ turns significantly negative around 2000.

when $h > 1$. Instead of computing [Newey and West \(1987\)](#) standard errors with a lag truncation parameter of $\lceil 1.5 \times h \rceil$, we find that one would draw almost identical inferences using [Hansen and Hodrick \(1980\)](#) standard errors with a lag truncation parameter equal to h . Going further, we show the estimates and our inferences are similar if we simply use non-overlapping h -month changes.

Finally, one might wonder if our dating of this break is due to distortions stemming from the 2009–2015 period when overnight nominal rates were stuck at the zero lower bound. Our use of 1-year rates as the independent variables in equations (1) and (2) limits any distortions since 1-year yields continued to fluctuate from 2009 to 2015. Indeed, even if we end our sample in 2007 or 2008, we still detect a break around 2000. For instance, if the post-2000 sample ends in December 2008, we find a daily $\beta_{day} = 0.77$ and a yearly $\beta_{12} = 0.20$, which are essentially indistinguishable from the numbers in Table I.

U.S. evidence prior to the Great Inflation. A natural explanation for the strong sensitivity of long-term nominal rates during the 1971-1999 sample is that this was a period when long-run inflation expectations became unanchored and were being continuously revised ([Gürkaynak et al., 2005](#)). Since inflation expectations have become firmly moored in recent decades, it is useful to compare the post-2000 patterns to the those pre-dating the Great Inflation, which ran from the late 1960s to the early 1980s. In the Internet Appendix, we examine the sensitivity of long-term Treasury yields to changes in short-term yields from 1953 (when the data become available) to 1968 (when inflation expectations began to drift up). Consistent with the view that inflation expectations were better-anchored prior to the Great Inflation, the 1953-1968 coefficients are lower than the 1971-1999 coefficients. However, while the level of β_h coefficients is lower in the 1953-1968 sample, we do not see the strong dependence on horizon h that is so evident in the post-2000 data. In summary, while the unanchoring and reanchoring of long-run inflation expectations may help explain shifts in the level of β_h over time, the strongly *frequency-dependent* sensitivity of long-term rates that we see since 2000 appears to be something new under the sun.

International evidence. Our focus is on the U.S., but it is useful to consider whether these same patterns are also observed in other large, highly-developed economies. In the Internet Appendix, we report estimates of equation (1) for the U.K., Germany, and Canada for both pre-2000 and post-2000 samples. Our data for Canada, Germany, and the U.K. begin in 1986, 1972, and 1985, respectively. We find similar patterns for Canada, Germany, and the U.K. Specifically, for all three countries, β_h is strongly decreasing in h in the post-2000 data, but not in the pre-2000 data.

III. YIELD-CURVE DYNAMICS AND BOND RETURN PREDICTABILITY

We now pinpoint the term-structure dynamics that account for the stronger high-frequency sensitivity and weaker low-frequency sensitivity of long rates that we see in the post-2000 data. We

show that this frequency-dependent sensitivity arises because, all else equal, past increases in short rates predict a subsequent flattening of the yield curve, as well as a subsequent decline in long-term yields and forwards. Statistically, this means that post-2000 yield curve dynamics are “path-dependent” or “non-Markovian”: it is not enough to know the current shape of the yield curve; instead, to form the best forecast of future bond yields and returns, one also needs to know how the yield curve has shifted in recent months. These non-Markovian dynamics are themselves a reflection of a new short-lived form of bond return predictability. Since 2000, the expected excess returns on long-term bonds are temporarily elevated following past increases in short rates. Thus, relative to an expectations-hypothesis baseline, long-term yields exhibit excess sensitivity at high frequencies and *temporarily overreact* to changes in short rates.

III.A. Non-Markovian yield-curve dynamics

We first show that strong horizon-dependence of β_h in the post-2000 period arises because yield curve dynamics have become non-Markovian.

Predicting level and slope. When examining term structure dynamics, it is useful to study the dynamics of yield-curve factors, especially level and slope factors (Litterman and Scheinkman, 1991). We define the level factor as the 1-year yield ($L_t \equiv y_t^{(1)}$) and the slope factor as the 10-year yield less the 1-year yield ($S_t \equiv y_t^{(10)} - y_t^{(1)}$)—a.k.a., the “term spread.” Most term structure models are Markovian with respect to current yield curve factors, meaning that the conditional mean of future yields depends only on today’s yield-curve factors. However, our key finding—the post-2000 horizon-dependence of the relationship between long- and short-term yields—suggests that it may be useful to include lagged factors when forecasting yields. This idea has proven useful in other contexts, including in Cochrane and Piazzesi (2005) and Duffee (2013). Specifically, we consider the following system of predictive monthly regressions:

$$(3a) \quad L_{t+1} = \delta_{0L} + \delta_{1L}L_t + \delta_{2L}S_t + \delta_{3L}(L_t - L_{t-6}) + \delta_{4L}(S_t - S_{t-6}) + \varepsilon_{L,t+1}$$

$$(3b) \quad S_{t+1} = \delta_{0S} + \delta_{1S}L_t + \delta_{2S}S_t + \delta_{3S}(L_t - L_{t-6}) + \delta_{4S}(S_t - S_{t-6}) + \varepsilon_{S,t+1}.$$

These regressions include level and slope as well as their changes over the prior six months, which is a simple way of allowing for longer lags without estimating too many free parameters.

Table II reports estimates of equations (3a) and (3b) for both the pre-2000 and post-2000 samples. We include specifications omitting all lagged changes, omitting lagged changes in slope, and including all predictors. Based on the Akaike information criterion (AIC) or Bayesian information criterion (BIC), the model in column (1) with no lagged changes is chosen in the pre-2000 sample, while the model in column (5) with lagged changes in level is selected in the post-2000 sample. In

the post-2000 sample, the lagged change in level is a highly significant negative predictor of the future slope—i.e., increases in the level of yields predict subsequent yield-curve flattening. For example, as shown in column (5), a 100 basis point increase in the level over the prior 6-months is associated with a 11 basis *per-month* decline in slope in the post-2000 sample ($p\text{-val} < 0.001$). By contrast, as shown in column (2), the coefficient on $L_t - L_{t-6}$ in the pre-2000 sample is zero. And, we can easily reject the hypothesis that the coefficients on $L_t - L_{t-6}$ in the pre- and post-2000 samples are equal ($p\text{-val} < 0.001$).

The model in equations (3a) and (3b) can match the puzzling post-2000 horizon-dependent behavior of β_h that we documented above. This model can be written as a restricted vector autoregression (VAR) in $\mathbf{y}_t = (L_t, S_t)'$ of the form: $\mathbf{y}_{t+1} = \boldsymbol{\mu} + \mathbf{A}_1\mathbf{y}_t + \mathbf{A}_2\mathbf{y}_{t-6} + \boldsymbol{\varepsilon}_{t+1}$. Let $\Gamma_{ij}(h)$ denote the ij th element of the autocovariance of \mathbf{y}_t at a lag of h months—i.e., the ij th element of $\boldsymbol{\Gamma}(h) = E[(\mathbf{y}_t - E[\mathbf{y}_t])(\mathbf{y}_{t-h} - E[\mathbf{y}_{t-h}])']$. Given the estimated parameters from equations (3a) and (3b), we can work out $\Gamma_{ij}(h)$ to obtain the VAR-implied values of β_h in equation (1):

$$(4) \quad \beta_h = \frac{\text{Var}(L_t - L_{t-h}) + \text{Cov}(S_t - S_{t-h}, L_t - L_{t-h})}{\text{Var}(L_t - L_{t-h})} = 1 + \frac{2\Gamma_{12}(0) - \Gamma_{12}(h) - \Gamma_{12}(-h)}{2(\Gamma_{11}(0) - \Gamma_{11}(h))}.$$

In the pre-2000 sample, Table I reported estimates of $\beta_1 = 0.46$ and $\beta_{12} = 0.56$. In the post-2000 sample, the estimates are $\beta_1 = 0.66$ and $\beta_{12} = 0.23$. Table II reports the VAR-implied values of β_1 and β_{12} from equation (4). In the pre-2000 data, all of the VAR models can roughly match both β_1 and β_{12} . In the post-2000 sample, all models can match β_1 , but only the models that include lagged changes in level—i.e., models that allow for non-Markovian dynamics—can match the sharp decline in β_{12} . Specifically, if the post-2000 VAR does not include lagged changes as in column (4), the VAR-implied values of β_{12} would be 0.59 and would be nowhere near what we observe in the data.

Predictable reversals in long-term rates. These post-2000 non-Markovian dynamics imply that there are predictable reversals in long rates following past increases in short rates. In Table III we estimate specifications that are reminiscent of the Jorda (2005) “local projection” approach to estimating impulse-response functions. Specifically, we predict the future changes in 10-year yields and forwards from month t to $t+h$ using the current level (L_t) and slope (S_t) of the yield curve as well as the prior month’s change in level ($L_t - L_{t-1}$) and slope ($S_t - S_{t-1}$):

$$(5) \quad z_{t+h} - z_t = \delta_0^{(h)} + \delta_1^{(h)}L_t + \delta_2^{(h)}S_t + \delta_3^{(h)}(L_t - L_{t-1}) + \delta_4^{(h)}(S_t - S_{t-1}) + \boldsymbol{\varepsilon}_{t \rightarrow t+h}.$$

Table III reports estimates of equation (5) for $z_t = y_t^{(10)}$ and $f_t^{(10)}$ in the pre- and post-2000 samples for $h = 3, 6, 9,$ and 12 - month changes. In Figure IV, we plot the coefficients $\delta_3^{(h)}$ on $L_t - L_{t-1}$ for $h = 1, 2, \dots, 12$, tracing out the expected future change in z_t from month t to $t+h$ in response to

an unexpected change in the level of rates between $t - 1$ and t .

In the post-2000 data, there are predictable reversals in both 10-year yields and forwards following an increase in short rates. However, there is no such reversal in the pre-2000 data. For 10-year yields, Table III reports that $\delta_3^{(6)} = -0.36$ ($p\text{-val} = 0.07$) after $h = 6$ -months in the post-2000 data. (The difference between $\delta_3^{(6)}$ in the pre- and post-2000 data is significant with a p -value of 0.04.) Table I showed that, since 2000, a 100 bps increase in short-rates in month t is associated with a 66 bps contemporaneous rise of long-term yields. Thus, Table III suggests that 36 bps—or more than half—of this initial response is expected to reverse within 6 months. As in Table I, the post-2000 reversion in 10-year forwards is even larger in magnitude and is statistically stronger. For 10-year forwards, we have $\delta_3^{(6)} = -0.52$ ($p\text{-val} < 0.01$) and the difference between $\delta_3^{(6)}$ in the pre- and post-2000 data is highly significant ($p\text{-val} < 0.01$). In summary, Table III and Figure IV show that long rates appear to temporarily overreact to changes in short rates in the post-2000 data, but there was no such tendency before 2000.

To better understand these results, we decompose 10-year yields into the sum of a level component and a slope component as in Table II—i.e., $y_t^{(10)} = L_t + S_t$ —and plot the coefficients $\delta_3^{(h)}$ versus h for both level ($z_t = L_t$) and slope ($z_t = S_t$). Consistent with Table II, Table III shows that the predictable reversals in long-term yields reflects the juxtaposition of two opposing forces in the post-2000. First, past increases in short-term rates predict subsequent increases in short-term rates in the post-2000 data, perhaps owing to the Fed’s growing desire to gradually adjust short rates (Stein and Sunderam, 2018). However, past increase in short rates strongly predict a subsequent flattening of the yield curve since 2000. Since the latter effect outweighs the former, we see predictable reversals in long-term yields post-2000.

III.B. Predicting bond returns

We show that our main finding—the fact that, in recent years, β_h declines rapidly as a function of horizon h —reflects a new form of bond return predictability. Namely, this result arises because past increases in the level of rates lead to *temporary* rise in the expected excess returns on long-term bonds.

Results for 10-year bonds. The k -month log excess return on 10-year bonds over the riskless return on k -month bills, $(k/12)y_t^{(k/12)}$, is:

$$(6) \quad rx_{t \rightarrow t+k}^{(10)} \equiv (k/12)(y_t^{(10)} - y_t^{(k/12)}) - (10 - k/12)(y_{t+k}^{(10-k/12)} - y_t^{(10)}).$$

We forecast the k -month excess return on 10-year zero-coupon bonds using level, slope, and the 6-month past changes in these two yield-curve factors:

$$(7) \quad rx_{t \rightarrow t+k}^{(10)} = \delta_0 + \delta_1 L_t + \delta_2 S_t + \delta_3 (L_t - L_{t-6}) + \delta_4 (S_t - S_{t-6}) + \varepsilon_{t \rightarrow t+k}.$$

In Table IV, we report the results from estimating these predictive regressions for $k = 1, 3,$ and 6-month returns. Panel A reports the results for the pre-2000 sample and Panel B shows the post-2000 results.

In the post-2000 data, Table IV shows the past change in the level of rates is a robust predictor of the excess returns on long-term bonds. However, there is no such predictability in the pre-2000 data. For instance, in column (6) of Panel B, we see that, all else equal, a 100 bps increase in short-term rates over the prior 6 months is associated with a $\delta_3 = 166$ bps ($p\text{-val} < 0.01$) increase in expected 3-month bond returns and the difference between δ_3 in the pre- and post-2000 data is statistically significant ($p\text{-val} < 0.01$). In untabulated results, we find that the post-2000 return predictability associated with past increases in the level of rates is short-lived and generally dissipates after $k = 6$ months. In other words, past increases in the level of rates lead to a temporary increase in the risk premia on long-term bonds.⁴

Results for other bond maturities. In the Internet Appendix, we examine the predictability for bond maturities other than $n = 10$ years. If, as we argue, past increases in short rates temporarily raise the compensation that investors require for bearing interest-rate risk, this should have a larger impact on the expected returns of long-term bonds than intermediate bonds. However, such a short-lived increase in the compensation for bearing interest rate risk should have relatively constant or even a hump-shaped effect on the yield and forward curves. The intuition is that the impact on bond yields equals the effect on a bond’s average expected returns over its lifetime. As a result, a *temporary* rise in the compensation for bearing interest rate risk can have a greater impact intermediate-term yields than on long-term yields. Indeed, this is precisely what we find in the post-2000 data.

In summary, since 2000, term premia on long-term bonds are *temporarily* elevated following increases in short rates. This implies that, relative to an expectations-hypothesis baseline, long rates *temporarily overreact* to movements in short rates and exhibit “excess sensitivity” at high frequencies.

III.C. Interpreting the evidence

Before developing our economic modelling framework, we pause to interpret our results. Our findings all point towards the view that, in recent years, the term premium on long-term bonds is increasing in the recent *change* in short rates, all else equal. This simple non-Markovian assumption can match the facts that, in the post-2000 data, (i) the sensitivity of long rates β_h declines with horizon h and (ii) that, controlling for current yield curve factors, past changes in short rates

⁴Consistent with the predictable curve flattening discussed above, the Internet Appendix shows that the predictability we find for 10-year bond returns is related to predictability for portfolios that locally mimic changes in the slope factor.

predict future yield-curve flattening, declines in long rates, and high excess returns on long-term bonds.

To develop these ideas, we shift notation slightly. Rather than identifying specific maturities, we now refer to the long-term yield as y_t and the short rate as i_t . We split the long-term yield into an expectations-hypothesis component, eh_t , that reflects expected future short-term rates and a term premium component, tp_t , that reflects expected future bond risk premia: $y_t = eh_t + tp_t$. Thus, by definition, β_h —the total sensitivity of long-term yields at horizon h —is the sum of the expectations-hypothesis, β_h^{eh} and the term premium, β_h^{tp} , components:

$$(8) \quad \frac{\overbrace{\text{Cov}[y_{t+h} - y_t, i_{t+h} - i_t]}^{\beta_h}}{\text{Var}[i_{t+h} - i_t]} = \frac{\overbrace{\text{Cov}[eh_{t+h} - eh_t, i_{t+h} - i_t]}^{\beta_h^{eh}}}{\text{Var}[i_{t+h} - i_t]} + \frac{\overbrace{\text{Cov}[tp_{t+h} - tp_t, i_{t+h} - i_t]}^{\beta_h^{tp}}}{\text{Var}[i_{t+h} - i_t]}.$$

First, consider the expectations-hypothesis piece. For now, assume the short-rate follows a univariate AR(1) process, implying $eh_t = \alpha^{eh} + \beta^{eh} \cdot i_t$ and $\beta_h^{eh} = \beta^{eh}$ for all h . Next, consider the term premium piece. In conventional asset-pricing theories, term premia only vary at business-cycle frequencies, so one would not expect β_h^{tp} to vary strongly with monthly horizon h . Thus, conventional theories suggest that $tp_t \approx \alpha^{tp} + \beta^{tp} \cdot i_t$, implying that $\beta_h^{tp} = \beta^{tp}$ and $\beta_h = (\beta^{eh} + \beta^{tp})$ for all h . In other words, it is difficult for conventional theories to match the strong horizon-dependence of β_h seen in the post-2000 data.

To generate horizon-dependent sensitivity, consider, instead, the following non-Markovian assumption:

$$(9) \quad tp_t = \alpha^{tp} + \beta^{tp} \cdot i_t + \delta^{tp} \cdot (i_t - i_{t-1}),$$

where $\delta^{tp} > 0$. This assumption implies that term premia depend on the current level of short rates and the recent *change* in short rates. Under this assumption, one can show that:

$$(10) \quad \beta_h = \beta^{eh} + \beta^{tp} + \delta^{tp} \cdot (1 - \gamma_h) \text{ where } \gamma_h \equiv \frac{\text{Cov}[i_{t+h-1} - i_{t-1}, i_{t+h} - i_t]}{\text{Var}[i_{t+h} - i_t]}.$$

The key is then to note that γ_h —the coefficient from a regression of $(i_{t+h-1} - i_{t-1})$ on $(i_{t+h} - i_t)$ —is an increasing function of h . When $\delta^{tp} > 0$, this in turn explains why β_h^{tp} is decreasing in h .⁵ Furthermore, when $\delta^{tp} > 0$, controlling for current level of short rates, the past change in short rates predicts future yield curve flattening, declines in long-term yields, and high excess returns on long-term bonds.

⁵For instance, if i_t follows an AR(1) of the form $i_{t+1} - \bar{i} = \rho_i(i_t - \bar{i}) + \varepsilon_{i,t+1}$, then $\gamma_h = (2\rho_i - \rho_i^{h-1} - \rho_i^{h+1}) / (2 - 2\rho_i^h)$. We have $\gamma_1 = -(1 - \rho_i)^2 / (2 - 2\rho_i) < 0$ and $\lim_{h \rightarrow \infty} \gamma_h = \rho_i > 0$. And, treating γ_h as continuous in h , we have $\partial \gamma_h / \partial h > 0$.

IV. A MODEL OF TEMPORARY BOND MARKET OVERREACTION

Because our key finding reflects a form of short-lived return predictability, the most natural explanations involve temporary supply-and-demand imbalances in financial markets (De Long et al. (1990), Shleifer and Vishny (1997), and Duffie (2010)). Our model emphasizes what we call “rate-amplifying” shocks to the supply and demand for long-term bonds. In Section V., we build on this framework and explicitly microfound three rate-amplification mechanisms—mortgage refinancing waves, investor extrapolation, and investor reaching-for-yield—and assess empirically the extent to which each mechanism helps explain our key finding. The model here emphasizes the common underlying structure and shared asset-pricing implications of these rate-amplification mechanisms. By contrast, Section V. emphasizes the idea that different amplification mechanisms have implications for different financial quantities.

Model setting. Time is discrete and infinite. Risk-averse bond arbitrageurs can hold either risky long-term nominal bonds or riskless short-term nominal bonds. The interest rate on short-term bonds follows an exogenous stochastic process. Long-term bonds are available in a given net supply that must be absorbed by bond arbitrageurs. Since the risk-bearing capacity of these specialized bond arbitrageurs is limited, shifts in the net supply of long-term bonds impact the term premium component of long-term yields as in Greenwood and Vayanos (2014) and Vayanos and Vila (2020).

We add two novel ingredients to this familiar setup. First, there are rate-amplifying supply-and-demand shocks: shocks to the net supply of long-term bonds are *positively* correlated with shocks to short rates. To induce arbitrageurs to absorb these supply shocks, the term premium component of long yields must increase when short rates rise, generating “excess sensitivity” of long-term yields relative to the expectations hypothesis baseline.

Second, arbitrage capital is slow-moving as in Duffie (2010): these net supply shocks walk down a short-run demand curve that is steeper than the long-run demand curve.⁶ Thus, an increase in short rates leads to a *temporary* supply-and-demand imbalance in the market for long-term bonds and, thus, a short-lived increase in bond risk premia. As a result, the excess sensitivity of long rates is greatest at short horizons. Furthermore, this frequency-dependent excess sensitivity is most pronounced when the underlying rate-amplifying net supply shocks are themselves transitory.

The model can match our key finding in Section II.—that β_h has fallen for large h and risen for small h post-2000—if (i) shocks to short-term nominal rates have become less persistent and (ii) the kinds of rate-amplification mechanisms we emphasize have grown in importance. We argue that (i) is justified by the strong evidence that shocks to the persistent component of inflation have become less volatile since the mid-1990s (Stock and Watson, 2007). We argue that (ii) is justified

⁶Our model is related to Greenwood et al. (2018), who incorporate slow-moving capital into a model of the term structure.

since these rate amplification mechanisms appear to have become more powerful over time.

Short- and long-term nominal bonds. At time t , investors learn that short-term bonds will earn a riskless log return of i_t in nominal terms between time t and $t + 1$. Short-term nominal bonds are available in perfectly elastic supply at this interest rate. One can think of the short-term nominal interest rate as being determined outside the model by monetary policy.

Long-term nominal bonds are available in a given net supply s_t that must be absorbed by the arbitrageurs in our model. The long-term nominal bond is a perpetuity. To generate a tractable linear model, we use the well-known [Campbell and Shiller \(1988\)](#) log-linear approximation to the return on this perpetuity. The log excess return on long-term bonds over short-term bonds from t to $t + 1$ is approximately:

$$(11) \quad rx_{t+1} \equiv \ln(1 + R_{t+1}) - i_t \approx \frac{1}{1 - \phi} y_t - \frac{\phi}{1 - \phi} y_{t+1} - i_t,$$

where y_t is the log yield-to-maturity on long-term bonds, $\phi \in (0, 1)$, and $D = 1/(1 - \phi)$ is the bond's duration—i.e., the sensitivity of the bond's price to its yield. Iterating equation (11) forward and taking expectations, the yield on long-term bonds is:

$$(12) \quad y_t = \overbrace{(1 - \phi) \sum_{j=0}^{\infty} \phi^j E_t [i_{t+j}]}^{eh_t} + \overbrace{(1 - \phi) \sum_{j=0}^{\infty} \phi^j E_t [rx_{t+j+1}]}^{tp_t}.$$

The long yield is the sum of an expectations hypothesis piece, eh_t , that reflects expected future short rates and a term premium, tp_t , reflecting expected future excess returns on long bonds over short bonds.

Arbitrageurs. There are two groups of specialized bond arbitrageurs, each with identical risk tolerance τ , who differ solely in the frequency with which they can rebalance their bond portfolios.

The first group of arbitrageurs are “fast-moving” and are free to adjust their holdings of long-term and short-term bonds each period. Fast-moving arbitrageurs are present in mass q and we denote their demand for long-term bonds at time t by b_t . Fast-moving arbitrageurs have mean-variance preferences over 1-period portfolio log returns. Their demand for long-term bonds at time t is:

$$(13) \quad b_t = \tau \frac{E_t [rx_{t+1}]}{\text{Var}_t [rx_{t+1}]}.$$

The second group of arbitrageurs are “slow-moving” and can only rebalance their holdings of long-term and short-term bonds every $k \geq 2$ periods. Slow-moving arbitrageurs are present in mass $1 - q$. A fraction $1/k$ of slow-moving arbitrageurs are active each period and can rebalance their portfolios, but then cannot trade again for the next k periods. As in [Duffie \(2010\)](#), this is a reduced-

form way to model the forces, whether due to institutional frictions or limited attention, that limit the speed of arbitrage capital flows. Since they only rebalance every k periods, slow-moving arbitrageurs have mean-variance preferences over their k -period *cumulative* portfolio excess return. Thus, the demand for long-term bonds from the subset of slow-moving arbitrageurs who are active at time t is:

$$(14) \quad d_t = \tau \frac{E_t[\sum_{j=1}^k rx_{t+j}]}{\text{Var}_t[\sum_{j=1}^k rx_{t+j}]}.$$

Holders of long-term bonds face two different types of risk. First, they are exposed to *short rate risk*. they will suffer a capital loss on their long-term bond holdings if short-term rates unexpectedly rise. Second, they are exposed to *supply risk*: there are shocks to the net supply of long-term bonds that impact the term premium component of long-term bond yields. We make the following assumptions about the evolution of these two risk factors.

Short-term nominal interest rates. The short-term nominal interest rate is the sum of a highly persistent component $i_{P,t}$ and a more transient component $i_{T,t}$:

$$(15) \quad i_t = i_{P,t} + i_{T,t}.$$

A natural interpretation is that the persistent component reflects long-run inflation expectations and the transient component reflects cyclical variation in short-term real rates and expected inflation. The persistent component $i_{P,t}$ follows an exogenous AR(1) process:

$$(16) \quad i_{P,t+1} = \bar{i} + \rho_P (i_{P,t} - \bar{i}) + \varepsilon_{P,t+1},$$

where $0 < \rho_P < 1$ and $\text{Var}_t[\varepsilon_{P,t+1}] = \sigma_P^2$. The transient component $i_{T,t}$ also follows an exogenous AR(1):

$$(17) \quad i_{T,t+1} = \rho_T i_{T,t} + \varepsilon_{T,t+1},$$

where $0 < \rho_T \leq \rho_P < 1$ and $\text{Var}_t[\varepsilon_{T,t+1}] = \sigma_T^2$.

If $\rho_T < \rho_P$ and σ_P is large relative to σ_T , then short-term nominal rates will be highly persistent. As a result, long-term nominal rates will be highly sensitive to movements in short-term nominal rates due to standard expectations-hypothesis logic. Indeed, a large value of σ_P is a good explanation for the high sensitivity of long-term rates observed in the 1970s, 1980s, and the 1990s when long-run inflation expectations were not well-anchored (Gürkaynak et al., 2005). However, long-run inflation expectations have become firmly anchored in recent decades and there is strong evidence that shocks to the persistent component of nominal inflation have become far less volatile since the mid-1990s.

Rate-amplifying shocks to the net supply of long-term bonds. Long-term nominal bonds are available in an exogenous, time-varying net supply s_t that must be held in equilibrium by arbitrageurs. This net supply equals the gross supply of long-term bonds minus the demand from other, non-arbitrageur investors outside the model who have inelastic demands. We assume that s_t follows an AR(1) process:

$$(18) \quad s_{t+1} = \bar{s} + \rho_s (s_t - \bar{s}) + C\varepsilon_{P,t+1} + C\varepsilon_{T,t+1} + \varepsilon_{s,t+1},$$

where $0 < \rho_s \leq \rho_T < 1$, $C \geq 0$, and $\text{Var}_t [\varepsilon_{s,t+1}] = \sigma_s^2$.

When $C > 0$, there are *rate-amplifying* net supply shocks—shocks to short rates are positively associated with shocks to net bond supply—and C parameterizes the strength of these amplification mechanisms. Equation (18) is a reduced-form way of capturing three different rate-amplification mechanisms that we detail in Section V.: (i) mortgage refinancing waves, (ii) investors who extrapolate recent changes in short rates, and (iii) investors who “reach for yield” by buying more long-term bonds when short rates are low. Rate-amplifying net supply shocks can arise either because increases in short rates are associated with increases in the gross supply of long-term bonds (as in the mortgage refinancing channel) or because they are associated with reductions in the demands of other, non-arbitrageur investors (as in the investor extrapolation and reaching-for-yield channels). The $\varepsilon_{s,t+1}$ shocks in (18) capture forces that are unrelated to short rates which also impact the net supply of long-term bonds. While the model can be solved for any arbitrary correlation structure between the $\varepsilon_{P,t+1}$, $\varepsilon_{T,t+1}$, and $\varepsilon_{s,t+1}$ shocks, we assume, for simplicity, that these three shocks are mutually orthogonal.

The difference between the persistence of these rate-amplifying net supply shocks and that of the underlying shocks to short-term rates plays an important role in our model’s ability to generate excess sensitivity that is most pronounced at high frequencies. To see why, note that equation (18) implies that the net supply of long-term bond is given by

$$(19) \quad s_t = \bar{s} + C[(i_{P,t} - \bar{i}) - (\rho_P - \rho_s) \sum_{j=0}^{\infty} \rho_s^j (i_{P,t-j-1} - \bar{i})] \\ + C[i_{T,t} - (\rho_T - \rho_s) \sum_{j=0}^{\infty} \rho_s^j i_{T,t-j-1}] + [\sum_{j=0}^{\infty} \rho_s^j \varepsilon_{s,t-j}].$$

When $\rho_s < \rho_T$, the rate-amplifying net supply shocks are less persistent than the underlying short rate shocks. As a result, net bond supply is increasing in the differences between the current level of each component of the short rate and a geometric moving-average of its past values. Thus, when $\rho_s < \rho_T$, s_t will be high when short rates have recently risen. By contrast, if $\rho_s = \rho_T = \rho_P$, s_t will be just as persistent as short rates. In this case, $s_t = \bar{s} + C(i_t - \bar{i}) + [\sum_{j=0}^{\infty} \rho_s^j \varepsilon_{s,t-j}]$ and only the current *level* of short rates—as opposed to recent changes in short rates—impacts net bond supply.

Equilibrium yields. At time t , there is a mass q of fast-moving arbitrageurs, each with demand b_t , and a mass $(1 - q)k^{-1}$ of active slow-moving arbitrageurs who rebalance their portfolios, each with demand d_t . These arbitrageurs must accommodate the *active net supply* of long-term bonds, which is the total net supply s_t less any supply held by inactive slow-moving arbitrageurs who do not rebalance at time t , $(1 - q)k^{-1} \sum_{j=1}^{k-1} d_{t-j}$. Thus, the market-clearing condition for long-term bonds at time t is:

$$(20) \quad \underbrace{qb_t}_{\text{Fast demand}} + \underbrace{(1 - q)k^{-1}d_t}_{\text{Active slow demand}} = \underbrace{s_t}_{\text{Total net supply}} - \underbrace{(1 - q)(k^{-1} \sum_{j=1}^{k-1} d_{t-j})}_{\text{Inactive slow holdings}}.$$

We conjecture that equilibrium yields y_t and the demands of active slow-moving arbitrageurs d_t are linear functions of a state vector, \mathbf{x}_t , that includes the steady-state deviations of both components of short-term nominal interest rates, the net supply of long-term bonds, and holdings of bonds by inactive slow-moving arbitrageurs. Formally, we conjecture that the yield on long-term bonds is $y_t = \alpha_0 + \alpha'_1 \mathbf{x}_t$ and that slow-moving arbitrageurs' demand for long-term bonds is $d_t = \delta_0 + \delta'_1 \mathbf{x}_t$, where the $(k + 2) \times 1$ dimensional state vector, \mathbf{x}_t , is given by $\mathbf{x}_t = [i_{P,t} - \bar{i}, i_{T,t}, s_t - \bar{s}, d_{t-1} - \delta_0, \dots, d_{t-(k-1)} - \delta_0]'$. These assumptions imply that the state vector follows a VAR(1) process $\mathbf{x}_{t+1} = \Gamma \mathbf{x}_t + \varepsilon_{t+1}$, where Γ depends on the parameters δ_1 governing slow-moving arbitrageurs' demand.

In the Internet Appendix, we show how to solve for equilibrium yields in this setting. A rational expectations equilibrium of our model is a fixed point of a specific operator involving the “price-impact” coefficients, (α'_1) , which show how the state variables impact bond yields, and the “demand-impact” coefficients, (δ'_1) , which show how these variables impact the demand of active slow-moving investors. Specifically, let $\omega = (\alpha'_1, \delta'_1)'$ and consider the operator $\mathbf{f}(\omega_0)$ which gives (i) the price-impact coefficients that will clear the market for long-term bonds and (ii) the demand-impact coefficients consistent with optimization on the part of active slow-moving investors when agents conjecture that $\omega = \omega_0$ at all future dates. A rational expectations equilibrium of our model is a fixed point $\omega^* = \mathbf{f}(\omega^*)$. Solving the model involves numerically finding a solution to a system of $2k$ non-linear equations in $2k$ unknowns.

An equilibrium solution only exists if arbitrageurs are sufficiently risk tolerant (i.e., for τ sufficiently large). When an equilibrium exists, there can be multiple equilibria. Equilibrium non-existence and multiplicity of this sort are common in overlapping-generations, rational-expectations models such as ours where risk-averse arbitrageurs with finite investment horizons trade an infinitely-lived asset that is subject to supply shocks. Different equilibria correspond to different self-fulfilling beliefs that arbitrageurs can hold about the price-impact of supply shocks and, hence, the risks of holding long-term bonds. However, we always find a unique equilibrium that is stable in the sense that equilibrium is robust to a small perturbation in arbitrageurs' beliefs regarding

the equilibrium that will prevail in the future. Consistent with the “correspondence principle” of Samuelson (1947), this unique stable equilibrium has comparative statics that accord with standard economic intuition. We focus on this unique stable equilibrium in our analysis. See Greenwood et al. (2018) for an extensive discussion of these issues.

The sensitivity of long-term yields. We now explain the factors that shape the sensitivity of long-term rates in our model and how this sensitivity depends on horizon. Consider the model-implied counterpart of the empirical regression coefficient in equation (1). In the model, the coefficient β_h from a regression of $y_{t+h} - y_t$ on $i_{t+h} - i_t$ is:

$$(21) \quad \beta_h = \frac{\text{Cov}[y_{t+h} - y_t, i_{t+h} - i_t]}{\text{Var}[i_{t+h} - i_t]} = \frac{\alpha'_1 (2\mathbf{V} - \mathbf{\Gamma}^h \mathbf{V} - \mathbf{V}(\mathbf{\Gamma}')^h) \mathbf{e}}{\mathbf{e}' (2\mathbf{V} - \mathbf{\Gamma}^h \mathbf{V} - \mathbf{V}(\mathbf{\Gamma}')^h) \mathbf{e}},$$

where $\mathbf{V} = \text{Var}[\mathbf{x}_t]$ denotes the variance of the state vector \mathbf{x}_t and \mathbf{e} denotes the $(k+2) \times 1$ vector with ones in the first and second positions and zeros elsewhere.⁷ We can then establish the following result:

Proposition 1. *The dependence of the coefficient β_h on time horizon h is governed by (i) the persistence ρ_x of the three shocks $x \in \{s, T, P\}$, (ii) the volatilities of the two short-rate shocks, σ_T and σ_P , (iii) the strength of the rate-amplification mechanisms C , and (iv) the degree to which capital is slow moving q .*

1. *When there are no rate-amplifying net supply shocks ($C = 0$), changes in term premia are unrelated to shifts in short rates and long-term yields do not exhibit excess sensitivity. Furthermore,*

(a) *if $\rho_T = \rho_P$, β_h is independent of h , σ_T , and σ_P .*

(b) *if $\rho_T < \rho_P$, β_h is increasing in h ; the level of β_h falls with σ_T and rises with σ_P for all h .*

2. *When there are rate-amplifying net supply shocks ($C > 0$), changes in term premia are positively correlated with changes in short rates and long-term yields exhibit excess sensitivity. Furthermore,*

(a) *if $\rho_s = \rho_T = \rho_P$, and all capital is fast-moving ($q = 1$), then β_h is independent of h ;*

(b) *if $\rho_s \leq \rho_T = \rho_P$ and either (i) supply shocks are transient ($\rho_s < \rho_T$) or (ii) capital is slow-moving ($q < 1$), then β_h is decreasing in h ;*

(c) *if $\rho_s \leq \rho_T < \rho_P$, β_h can be non-monotonic in h .*

Proof. See the Internet Appendix for all proofs. □

⁷To derive this expression, note that $y_{t+h} - y_t = \alpha'_1 (\mathbf{x}_{t+h} - \mathbf{x}_t)$ and $i_{t+h} - i_t = \mathbf{e}' (\mathbf{x}_{t+h} - \mathbf{x}_t)$. Since the state-vector \mathbf{x}_t follows a VAR(1) process $\mathbf{x}_{t+1} = \mathbf{\Gamma} \mathbf{x}_t + \boldsymbol{\varepsilon}_{t+1}$, we have $\text{Var}[\mathbf{x}_{t+h} - \mathbf{x}_t] = 2\mathbf{V} - \mathbf{\Gamma}^h \mathbf{V} - \mathbf{V}(\mathbf{\Gamma}')^h$ and the result follows.

When there are no rate-amplifying net supply shocks ($C = 0$), long rates are not excessively sensitive to short rates when judged relative to the expectations hypothesis. If short rates contain both a transient and a more persistent component ($\rho_T < \rho_P$), one should actually expect the β_h coefficients to increase in horizon h when $C = 0$. This effect arises since movements in the persistent short rate component are associated with larger movements in long rates by standard expectations-hypothesis logic and because the persistent component dominates changes in short rates at longer horizons. Furthermore, when $\rho_T < \rho_P$, the level of β_h for any horizon h depends on σ_T and σ_P . For instance, an increase in σ_P raises the fraction of total short-rate variation that is due to the persistent component. Since shocks to the persistent component of short rates have a larger impact on long rates, an increase in σ_P raises β_h for all h .

Rate-amplifying net supply shocks ($C > 0$) generate excess sensitivity. However, Part 2.(a) of Proposition 1 shows that rate-amplification ($C > 0$) need not generate *horizon-dependent* excess sensitivity—i.e., *temporary* overreaction of long rates when judged relative to the expectations hypothesis. To generate β_h coefficients that decline with h , Part 2.(b) clarifies that either (i) the rate-amplifying net supply shocks must be less persistent than the underlying short-rate shocks ($\rho_s < \rho_T$) or (ii) these rate-amplifying shocks must be met by a slow-moving arbitrage response ($q < 1$). Under either of these conditions, shifts in short rates trigger a short-lived supply-and-demand imbalance in the market for long-term bonds, leading long rates to *temporarily* overreact to short rates. In practice, we suspect that *both* transitory rate-amplifying net supply shocks and slow-moving capital play a role in explaining why β_h declines steeply with h in the recent data. Furthermore, these two mechanisms reinforce one another: it is easier to quantitatively match the steep decline in β_h as a function of h in calibrations that feature both elements.⁸

Model calibration. Our main findings are that β_h has risen at high frequencies (low h) but has fallen at low frequencies (high h) in recent decades, leading β_h to decline steeply with horizon h in the post-2000 data. Guided by Proposition 1, we discuss how to understand the changing sensitivity of long rates. We focus on the role of changes in the volatility of persistent short rate shocks (σ_P) and the strength of any rate-amplifying mechanisms (C). Our model can match the data if (1) shocks to the persistent component of short-term nominal rates have become less volatile in the post-2000 period (σ_P has fallen) and (2) the rate-amplifying supply-and-demand mechanisms we

⁸Technically, when capital is slow-moving ($q < 1$) and $\rho_s \leq \rho_T = \rho_P$, β_h is only guaranteed to be *locally* decreasing in h for $h \leq k$ —i.e., for horizons shorter than over which all slow-moving arbitrageurs will have rebalanced their portfolios. While we always have $\beta_h < \beta_{h-1}$ for $h \leq k$, we can have $\beta_h > \beta_{h-1}$ for $h > k$. However, even when there are local non-monotonicities, β_h is globally decreasing in the sense that $\lim_{h \rightarrow \infty} \beta_h < \beta_1$. What explains the potential for these local non-monotonicities? As in Duffie (2010), the gradual adjustment of slow-moving arbitrageurs can give rise to modest echo effects for $h > k$, generating a series of damping oscillations that converge to $\lim_{h \rightarrow \infty} \beta_h$. These oscillations arise because the slow-moving arbitrageurs who reallocate soon after a supply shock lands take large opportunistic positions. These positions temporarily reduce the active supply of long-term bonds and need to be re-absorbed in later periods.

emphasize have become more important in recent decades (C has risen).

We consider an illustrative calibration of the model in which each period is a month. We assume the following parameters were the same in the 1971-1999 and post-2000 periods:

- **Persistence:** $\rho_P = 0.995$, $\rho_T = 0.96$, and $\rho_S = 0.80$. These parameters imply that shocks to the persistent short rate component have a half-life of 11.5 years, shocks to the transient component have a half-life of 1.4 years, and shocks to the net supply of long bonds have a half-life of 3 months.
- **Slow-moving capital:** $q = 30\%$ and $k = 12$. Thus, $1 - q = 70\%$ of the arbitrageurs are slow-moving and only rebalance their bond portfolios every 12 months. These assumptions capture the idea that many large institutional investors only rebalance their portfolios annually.
- **Volatility of the transient component of short rates:** $\sigma_T^2 = 0.15\%$.
- **No independent net supply shocks:** $\sigma_S^2 = 0$. This assumption is without loss of generality.
- **Other parameters:** $\tau = 0.5$ and $\phi = 119/120$, so the duration of the perpetuity is 10 years—i.e., $D = 1/(1 - \phi) = 120$ months.

For the pre-2000 period, we assume:

- **A large persistent component of short rates:** $\sigma_P^2 = 0.15\%$. The implied volatility of the short rate is 4.12% which compares with a volatility of 1-year yields of 2.63% in the 1971-1999 sample.
- **No rate-amplifying net supply shocks:** $C = 0$.

By contrast, for the post-2000 period, we assume:

- **A small persistent component of short rates:** $\sigma_P^2 = 0.012\%$. The implied standard deviation of the short rate is 1.77% which is similar to the post-2000 volatility of 1-year yields of 1.85%.
- **Net supply shocks induced by short rate shocks:** $C = 0.55$. Thus, we assume a meaningful increase in the strength of rate-amplifying supply-and-demand mechanisms.

The first graph in Figure V plots the model-implied regression coefficients β_h from equation (21) against monthly horizon h for the pre- and post-2000 calibrations. In the pre-2000 calibration where σ_P is large and $C = 0$, β_h is high and largely independent of h . In fact, β_h rises gradually with h —as it does in the pre-2000 data—because the more persistent component of short rates dominates when changes are computed at longer horizons. By contrast, in the post-2000 calibration

where σ_P is smaller and C is large, β_h declines steeply with h . And, since σ_P is lower, β_h reaches a lower level for large h .

β_h declines with h in the post-2000 calibration because short rate shocks give rise to transient rate-amplifying shocks to the net supply of long-term bonds ($C > 0$ and $\rho_s < \rho_T$) that encounter a short-run demand curve that is steeper than the long-run demand curve due to slow-moving capital ($q < 1$), triggering short-lived market imbalances. The second graph in Figure V shows that β_h only declines moderately with h in our post-2000 calibration if we drop the assumption that arbitrage capital is slow-moving. Thus, transient rate-amplifying net supply shocks and slow-moving capital are both helpful for quantitatively matching the fact that β_h declines so steeply with h in the post-2000 data.

Another way of understanding the mechanism is to study the model-implied impulse response functions following a surprise increase to short rates. We carry out this exercise in the Internet Appendix. To summarize, an initial positive shock to short rates leads to a rise in term premia. Thus, relative to the expectations-hypothesis, long rates are excessively sensitive to short rates. However, the rise in term premia wears off quickly, explaining our key finding that β_h declines sharply with horizon h . Furthermore, the initial rise in short rates predicts future yield curve flattening and future reversals in long rates.

In addition to matching the fact that β_h declines steeply with h in the post-2000 period, the model can also match the related empirical findings documented above. First, the model is consistent with our return forecasting evidence: in the post-2000 calibration, bond risk premia $E_t[rx_{t+1}] = \tau^{-1}V^{(1)}b_t$ are elevated when short-term rates have recently risen. Intuitively, if rate-amplifying net supply shocks ($C > 0$) are either transient ($\rho_s < \rho_T$) or are met by a slow-moving arbitrage response ($q < 1$), then fast-moving arbitrageurs must bear greater interest-rate risk when short rates have recently risen— b_t will be higher—and they will require additional compensation for bearing this extra risk

Second, let $L_t = i_t$ and $S_t = y_t - i_t$ denote the model-implied level and slope factors. If we estimate equation (3b) in data simulated from the model, we find that past increases in the level of rates predict a flattening of the yield curve in the post-2000 calibration but not in the pre-2000 calibration. In the post-2000 calibration, past increases in the level of rates are associated with a higher current risk premium on long-term bonds. Since the risk premium is $E_t[rx_{t+1}] = S_t - \phi(1 - \phi)^{-1}(E_t[\Delta S_{t+1}] + E_t[\Delta L_{t+1}])$, all else equal, $E_t[\Delta S_{t+1}]$ is lower when short rates have recently risen.

V. RATE-AMPLIFICATION MECHANISMS

We now explore three rate-amplification mechanisms—mortgage refinancing, extrapolation, and reaching for yield—that may help explain why increases in short rates trigger temporary supply-

and-demand imbalances in the market for long-term bonds. For each mechanism, we show how it microfounds rate-amplifying net supply shocks like those we introduced in reduced-form in Section IV. and then embed it the modelling framework developed above.⁹ Next, we discuss why the strength of each channel may have increased in recent decades. Finally, by examining the relationship between bond yields and different financial quantities, we assess empirically the extent to which each channel contributes to the frequency-dependent sensitivity of long-term rates we observe after 2000.

V.A. *Mortgage refinancing*

Negative shocks to short-term rates trigger mortgage refinancing waves in the U.S. that lead to temporary reductions in the effective gross supply of long-term bonds and, thus, temporary declines in bond term premia (Hanson, 2014; Malkhozov et al., 2016). The mortgage refinancing channel is only relevant in countries such as the U.S. where fixed-rate mortgages with an embedded prepayment option are an important source of mortgage financing. However, Domanski et al. (2017) point to a similar rate-amplification mechanism—one that may be more important in the Eurozone—stemming from the desire of insurers and pensions to dynamically match the duration of their assets and liabilities.

Modeling the mechanism. Most fixed-rate, residential mortgages in the U.S. give the borrower the option to prepay at any time without a penalty (Boyarchenko et al., 2019). When rates fall, the option to prepay and refinance older, higher-coupon mortgages becomes more attractive to borrowers. Households exercise their prepayment options only gradually after a decline in rates, leading the effective maturity or “duration” of outstanding mortgages—i.e., the sensitivity of mortgage prices to changes in interest rates—to decline when long-term rates fall. And, the amount of expected mortgage refinancing activity varies significantly over time: depending on the past path of rates, there are times when many households move from being far from refinancing to being close and vice versa. The resulting shifts in expected refinancing activity trigger large changes to the total quantity of interest rate risk that must be borne by bond market investors, leading to transient, but sizable fluctuations in bond term premia .

This mortgage refinancing channel can be used to micro-found a specification for the net supply of long-term bonds that is similar to equation (19). Following (Malkhozov et al., 2016), we assume that (i) there is a constant quantity M of outstanding fixed-rate mortgages with an embedded prepayment option; (ii) the primary mortgage rate, y_t^M , equals the long-term yield, y_t , plus a constant spread; (iii) the average coupon on outstanding mortgages evolves according to $c_{t+1}^M - c_t^M = -\eta \cdot (c_t^M - y_t^M)$, where $(c_t^M - y_t^M)$ is the “refinancing incentive” at time t and $\eta \in [0, 1]$ is the sensitivity of c_{t+1}^M to the refinancing incentive at t ; (iv) the average “duration” or effective

⁹The Internet Appendix provides additional details and illustrative calibrations of these three microfounded models.

maturity of outstanding mortgages is $DUR_t^M = \overline{DUR}^M - N \cdot (c_t^M - y_t^M)$, where $\overline{DUR}^M > 0$ and $N > 0$ is the “negative convexity” of the average mortgage; and (v) the effective gross supply of long-bonds at time t is $s_t = M \cdot DUR_t^M$.

Each of these assumptions captures a well-known and reliable empirical regularity about the U.S. mortgage market. In particular, assumption (iii) captures the fact that, when the refinancing incentive $(c_t^M - y_t^M)$ is higher, more households refinance their existing high-coupon mortgages at time t , leading the average coupon to fall from t to $t + 1$. Assumption (iv) captures the fact that, when the refinancing incentive is higher, more households are expected to refinance their existing mortgages in the near future, implying that the average outstanding mortgage behaves more like a short-term bond. These assumptions imply that the effective gross supply of long-term bonds at time t is:

$$(22) \quad s_t = M \cdot \overline{DUR}^M + MN \cdot (y_t - \eta \sum_{j=0}^{\infty} (1 - \eta)^j y_{t-1-j}).$$

Thus, the mortgage refinancing channel implies that bond investors must bear greater interest rate risk when long-term rates have recently risen. And, the strength of this channel is given by the product MN .

There have been two structural shifts that are relevant for the strength of the refinancing channel. First, mortgage-backed securities (MBS) have become a larger share of the U.S. bond market over time. In the language of the model, this means that M has risen. From 1976 to 1999, MBS on average accounted for 21% of the value of the Bloomberg-Barclays Aggregate Index, a proxy for the broad U.S. bond market. From 2000 to 2019, the corresponding figure was 33%. As a result, movements in the duration of the outstanding mortgages (DUR_t^M) now generate far larger shifts in the effective supply of long-term bonds when judged relative to the overall U.S. bond market. Second, due a secular decline in refinancing costs, mortgage refinancing has become more interest-rate elastic over time (Bennett and Peristiani, 2001; Fuster et al., 2019). More elastic refinancing corresponds to a rise in both N and η . As a result, the association between DUR_t^M and recent changes in long rates has grown stronger. Together these changes suggest that the strength of the mortgage refinancing channel has grown in recent decades.

To solve our model of mortgage refinancing, we substitute the expression for supply in (22) into the market-clearing condition in (20) from Section IV.. As above, fraction q of investors are fast-moving with demands given by equation (13) and fraction $(1 - q)$ are slow-moving and only rebalance their portfolios every $k \geq 2$ periods with demands given by (14). We can then establish the following proposition:

Proposition 2. Mortgage refinancing model. *For simplicity suppose $\rho_T = \rho_P$. When $MN > 0$, long-term yields are excessively sensitive to short rates. When $MN > 0$ and $\eta = 0$, this excess sensitivity is only horizon-dependent—i.e., the model-implied regression coefficients β_h in equation*

(21) only decline with horizon h —when arbitrage capital is slow moving ($q < 1$). By contrast, when $MN > 0$ and $\eta > 0$, β_h declines with horizon h even if all arbitrage capital is fast-moving ($q = 1$).

When $\eta > 0$, shocks to short rates trigger shifts in effective bond supply that are less persistent than the underlying short rate shocks, giving rise to horizon-dependent excess sensitivity even without slow-moving capital. However, we are best able to quantitatively match the post-2000 behavior of the β_h coefficients using calibrations of this model in which (i) MN has risen substantially from the pre-2000 level and (ii) the resulting rate-amplifying supply shocks are met by a gradual arbitrage response. In addition, our model of mortgage refinancing predicts that (1) mortgage duration DUR_t^M is high when interest rates have recently risen and (2) the level of mortgage duration positively predicts future excess returns on long-term bonds (i.e., $E_t [rx_{t+1}]$ is high when DUR_t^M is high).

Evidence from mortgage-related quantities. To assess the contribution of the refinancing channel to our findings, we use two proxies for the impact of mortgage refinancing on the effective supply long-term bonds. The first is $y_t^M - c_t^M$, the mortgage refinancing *disincentive*. Here y_t^M is the average primary rate for 30-year, fixed-rate mortgages from Freddie Mac’s Primary Mortgage Market Survey and c_t^M is the average outstanding coupon of MBS in the Bloomberg-Barclays U.S. MBS index. The index covers pass-through MBS backed by conventional fixed-rate mortgages that are guaranteed by Fannie Mae, Freddie Mac, and Ginnie Mae. This refinancing disincentive measure, which is associated with a higher duration on outstanding mortgages, is available beginning in Jan-1976. The second is the duration-to-worst of the Bloomberg-Barclays U.S. MBS index, DUR_t^M , a measure of the sensitivity of MBS prices to changes in long-term yields. This MBS duration measure is available on a monthly basis beginning in Jan-1976.¹⁰ The correlation between $y_t^M - c_t^M$ and DUR_t^M is 0.55 from 1976 to 1999 and 0.66 in the post-2000 sample.

Using each of these proxies (X_t) for mortgage duration, we first estimate

$$(23) \quad X_t = \gamma_0 + \gamma_1 L_t + \gamma_2 S_t + \gamma_3 (L_t - L_{t-6}) + \gamma_4 (S_t - S_{t-6}) + \varepsilon_t^{MBS},$$

for the pre- and post-2000 samples. We are interested in the coefficient on $L_t - L_{t-6}$, which tells us how MBS duration responds to recent changes in the level of short rates. Second, we estimate

$$(24) \quad rx_{t \rightarrow t+3}^{(10)} = \delta_0 + \delta_1 L_t + \delta_2 S_t + \delta_3 (L_t - L_{t-6}) + \delta_4 (S_t - S_{t-6}) + \delta_5 X_t + \varepsilon_{t \rightarrow t+3}^{(10)},$$

for the pre-2000 and post-2000 samples. That is, we run horse race regressions to assess whether

¹⁰Barclays uses its proprietary prepayment model to estimate the expected cashflows for each MBS. Yield-to-worst is the internal rate of return that equates MBS price and the present value of expected cash flows. Barclays computes the Macaulay duration of MBS treating these expected cashflows as given and index duration is the valued-weighted average of security-level durations. Beginning in 1989, Barclays reports the option-adjusted duration measure used in Hanson (2014). In the post-2000 sample, this slightly more sophisticated measure has a correlation of 0.84 with the measure we use.

mortgage waves help explain why past changes in short rates forecast excess bond returns in the post-2000 data. We are interested in the coefficients on X_t and $L_t - L_{t-6}$ and how these coefficients change when these two variables are included jointly as opposed to one at a time.

Panel A of Table V shows the results from estimating (23) using the refinancing disincentive ($X_t = y_t^M - c_t^M$) and shows that it is responsive to past changes in level. Comparing columns (3) and (6), we see that it has become more responsive to past changes in level since 2000 (p -value = 0.06). Panel B reports the results for estimating this same equation using the duration of the Barclays MBS index ($X_t = DUR_t^M$) and delivers a similar message. Panel C reports the results from estimating (24) using the refinancing disincentive and suggest that the refinancing channel helps explain why past changes in the level of rates predict high excess returns on long-term bonds in the post-2000 data. In Panel C column (5), we see that $y_t^M - c_t^M$ attracts a positive and significant coefficient when forecasting 3-month excess bond returns after 2000. By contrast, the corresponding coefficient in column (2) for the pre-2000 sample is near zero and insignificant. And, the difference between the coefficients in columns (2) and (5) is significant (p -value < 0.01). However, when we use both $(y_t^M - c_t^M)$ and $(L_t - L_{t-6})$ to forecast $rx_{t \rightarrow t+3}^{(10)}$ in column (6), the coefficients on both variables decline noticeably relative to those in columns (4) and (5) where they are considered in isolation. This is precisely what we should expect if the refinancing channel plays an important role in explaining the short-lived excess sensitivity that we see in the post-2000 data.¹¹ Panels D shows that DUR_t^M also strongly forecasts $rx_{t \rightarrow t+3}^{(10)}$ in the post-2000 sample with the expected signs.

V.B. Investor extrapolation

Several recent papers, including Cieslak (2018), Giglio and Kelly (2018), Brooks et al. (2019) and D’Arienzo (2020) argue that some bond investors make biased forecasts of future interest rates. Positive shocks to short rates lead extrapolative investors to overestimate the future path of short rates and demand fewer long-term bonds. As a result, the quantity of bonds that must be held by unbiased investors rises when short rates rise, leading to an rise in term premium and generating excess sensitivity. If these expectational errors are transitory or if the arbitrage response is slow, extrapolation will create a short-lived market imbalance, leading long rates to temporarily overreact to changes in short rates.

Modeling the mechanism. We assume that some investors have “diagnostic expectations” about short rates in the sense that they “overweight future outcomes that have become more likely in

¹¹If mortgage refinancing was the only source of rate amplification in the U.S. bond market and there was no slow-moving arbitrage capital, then $(y_t^M - c_t^M)$ should be a sufficient statistic for bond risk premia and should completely drive out $(L_t - L_{t-6})$ in a horse race specification. However, if mortgage refinancing was one of several amplification mechanisms, or if arbitrage capital was slow moving, then one expected both $(y_t^M - c_t^M)$ and $(L_t - L_{t-6})$ to attract meaningful coefficients.

light of incoming data” (Bordalo et al., 2017). In contrast to most recent work on diagnostic expectations—which takes a representative agent approach—we adopt a heterogenous agent approach, enabling us to study the dynamic arbitrage response of unbiased bond investors to these rate-amplifying demand shocks.

We assume that fraction f of investors have diagnostic expectations about short rates. Diagnostic investors’ demand for long-term bonds is $h_t = \tau(E_t^D[rx_{t+1}]/\text{Var}_t^D[rx_{t+1}])$ where $E_t^D[\cdot]$ denotes diagnostic investors’ biased expectations. Fraction $(1 - f)$ of a bond investors have fully rational expectations about short-term interest rates. Of these rational investors, fraction q are fast-moving with demands given by equation (13) and fraction $(1 - q)$ are slow-moving with demands given by (14). We assume the gross supply of long-term bonds is constant over time at $s_t = \bar{s}$. Following Maxted (2020), we assume that diagnostic investors’ expectation of the transient component of short-term rates ($i_{T,t}$) is

$$(25) \quad E_t^D[i_{T,t+1}] = \rho_T i_{T,t} + \theta \cdot [i_{T,t} - (\rho_T - \kappa_T) \sum_{j=0}^{\infty} \kappa_T^j i_{T,t-j-1}],$$

where $\theta \geq 0$ and $\kappa_T \in [0, \rho_T]$. The parameter θ governs the extent to which diagnostic expectations depart from rationality ($\theta = 0$) and κ_T governs the persistence of investors’ mistaken beliefs about short rates.¹² When $\theta > 0$ and $\kappa_T < \rho_T$, equation (25) says that diagnostic investors overestimate $i_{T,t+1}$ when $i_{T,t}$ has recently risen. Thus, extrapolation leads to a model that is similar to the reduced-form specification for net bond supply in equation (19). We adopt an analogous specification for diagnostic investors’ expectations of the persistent component of short rates ($i_{P,t}$), but assume for simplicity that diagnostic investors form rational forecasts of all other state variables.

The strength of this extrapolation channel is given by $f\theta$ —i.e., the mass of diagnostic investors (f) times the extent to which their expectations depart from rationality (θ). Why might $f\theta$ have risen in recent decades? While many bond investors may have a tendency to extrapolate past changes in interest rates, it is natural to think that this tendency is most pronounced amongst investors in bond mutual funds. Indeed, there is a long literature arguing that mutual fund investors—who are predominantly households and smaller institutions—tend to be more prone to common psychological biases than larger institutional investors (Barberis et al., 1998; Dichev, 2007; Frazzini and Lamont, 2008). Furthermore, mutual funds have become more important players in the U.S. bond market in recent decades. Based on data from Federal Reserve’s Financial Accounts, mutual funds’ share of Treasury and MBS holdings has gradually risen from 5% in the early 1990s to nearly 10% today. And, mutual funds have rapidly gained share in the corporate bond market, rising from a 7% share in early 2009 to over 20% today. Thus, even if individual

¹²In the limit where $\kappa_T = 0$, $E_t^D[i_{T,t+1}] = \rho_T i_{T,t} + \theta \varepsilon_{T,t}$, so investors’ mistakes ($\theta \varepsilon_{T,t}$) are serially uncorrelated over time. In the opposite limit where $\kappa_T = \rho_T$, $E_t^D[i_{T,t+1}] = \rho_T i_{T,t} + \theta i_{T,t}$, so investors’ mistakes ($\theta i_{T,t}$) are just as persistent as $i_{T,t}$.

mutual fund investors have not become more extrapolative since 2000 (i.e., if θ has not changed), this group of extrapolation-prone investors has become more important in the bond market (corresponding to a rise in f). In this setting, we can demonstrate the following result:

Proposition 3. Investor extrapolation model. *For simplicity suppose $\rho_T = \rho_P$ and $\kappa_T = \kappa_P$. When $f\theta > 0$, long rates are excessively sensitive to short rates. When $f\theta > 0$ and $\kappa_T = \rho_T$, this excess sensitivity is only horizon-dependent—i.e., the regression coefficients β_h only decline with horizon h —when unbiased arbitrage capital is slow moving ($q < 1$). By contrast, when $f\theta > 0$ and $\kappa_P < \rho_T$, β_h declines with horizon h even if all arbitrage capital is fast-moving ($q = 1$).*

When $f\theta > 0$ and $\kappa_P < \rho_P$, extrapolation generates transitory rate-amplifying demand shocks, giving rise to frequency-dependent excess sensitivity even without slow-moving capital. However, this frequency-dependent excess sensitivity becomes more pronounced when these demand shocks are met by a gradual arbitrage response from unbiased investors. Thus, we are best able to quantitatively match the post-2000 behavior of the β_h coefficients using calibrations of our extrapolation model in which (i) $f\theta$ has risen from its pre-2000 level and (ii) arbitrage is gradual. Our extrapolation model also predicts that: (1) the bond holdings of extrapolative investors, h_t , are low when interest rates have recently risen and (2) the level of extrapolative investors' bond holdings negatively predicts excess returns on long bonds.

Evidence from bond mutual fund flows. To assess whether investor extrapolation contributes to high-frequency excess sensitivity, we obtain monthly data from 1984 to 2019 on the total net assets of taxable bond mutual funds and the net dollar flows into these funds from the Investment Company Institute. We then compute the 3-month cumulative percentage flow into bond funds, $\%FLOW_{t-3 \rightarrow t}$. Using bond fund flows as a proxy for the rate-amplifying demand of extrapolative investors in our model (h_t), we first estimate equation (23) with $X_t = \%FLOW_{t-3 \rightarrow t}$ for the pre-2000 and post-2000 samples. The results are presented in Panel A of Table VI and show that bond mutual funds tend to suffer investor outflows when short-term interest rates decline. This result is consistent with the vast literature on return-chasing behavior by mutual fund investors (Warther, 1995; Sirri and Tufano, 1998). Interesting, this relationship is actually stronger in the pre-2000 sample than in the post-2000 sample, consistent with other evidence that mutual fund flows have become less performance sensitive in recent years. However, the importance of mutual funds within the bond market has increased meaningfully since 2000.

In Panel B, we estimate equation (24) with $X_t = \%FLOW_{t-3 \rightarrow t}$ for the pre- and post-2000 samples. As shown in column (5), past mutual fund flows predict low future excess returns on 10-year bonds in the post-2000 data. By contrast, as shown in column (2), there are no such relationships in the pre-2000 sample. However, when we use both $(L_t - L_{t-6})$ and $\%FLOW_{t-3 \rightarrow t}$ to forecast returns in column (6), the coefficients on both variables decline meaningfully relative to those shown in columns (4) and (5) where they are considered in isolation. As above, this is

what one would expect if extrapolation plays a role in explaining why long-term yields temporarily overreact to short rates in the post-2000 data.

V.C. Investors who reach for yield

Investors who “reach for yield” when short rates decline are a final potential source of rate-amplifying demand. According to the reaching-for-yield channel, negative shocks to short rates boost the demand for long-term bonds from “yield-seeking investors.” Holding fixed the gross supply, the net supply of long-term bonds that must be held by fast- and slow-moving arbitrageurs declines when short rates fall, leading term premia to decline when short rates fall.

Modeling the mechanism. We assume that fraction f of bond investors are “yield-seeking” and have non-standard preferences as in [Hanson and Stein \(2015\)](#). The idea is that, for either behavioral or institutional reasons, these investors only care about the current yield on their portfolios instead of their expected portfolio returns. Yield-seeking investors’ demand for long-term bonds is:

$$(26) \quad h_t = \tau \frac{y_t - i_t}{\text{Var}_t [rx_{t+1}]}.$$

Since $E_t [rx_{t+1}] = (y_t - i_t) - (\phi / (1 - \phi)) \cdot E_t [y_{t+1} - y_t]$, equation (26) means that yield-seeking investors neglect any expected capital gains or losses from holding long-term bonds. And, because expectations-hypothesis logic implies that long-term yields are expected to rise when short rates are low, this implies that these investors have an elevated demand for long bonds when short rates are low. As above, the gross supply of long-term bonds is constant. A mass $(1 - f)$ of a bond investors are expected-return-oriented and have standard mean-variance preferences. Of these, fraction q are fast-moving with demands given by (13) and fraction $(1 - q)$ are slow-moving with demands given by (14). And, prior research suggests the reaching-for-yield channel may have grown stronger—corresponding to a rise in f —in recent years as interest rates have reached historically low levels.¹³ Using this model, we can then show:

Proposition 4. *Investor reaching for yield model.* *Suppose $\rho_T = \rho_P$. When $f > 0$, long rates are excessively sensitive to short rates. However, this excess sensitivity is only horizon-dependent when arbitrage capital is slow moving ($q < 1$).*

Our model of reaching-for-yield also predicts that: (1) the bond holdings of yield-seeking investors, h_t , are low when interest rates are high and (2) the level of yield-seeking investors’ bond holdings negatively predicts future excess returns on long-term bonds.

¹³[Lian et al. \(2017\)](#) provide experimental evidence that the tendency to take on greater risk when short rates decline becomes more pronounced when the level of short rates is already low. Building on Prospect Theory ([Kahneman and Tversky, 1979](#)), they argue that this yield-seeking behavior becomes more pronounced as rates fall further below some reference level that investors are accustomed to based on past experience.

Since these rate-amplifying shifts in demand are tied to the level of short rates as opposed to recent changes in short rates, reaching for yield generates persistent shifts in demand. Thus, while reaching-for-yield can generate excess sensitivity, in the absence of gradual arbitrage, this excess sensitivity is not horizon-dependent. And, while the combination of reaching-for-yield and gradual arbitrage generates horizon-dependent sensitivity, our calibrations of this model struggle to quantitatively match the profile of β_h seen in the post-2000 data. Thus, it is not clear that reaching-for-yield can explain why excess sensitivity has become so pronounced at high frequencies. Going further, reaching-for-yield itself may be a slow-moving phenomenon: investors may only gradually take on greater risk following a decline in short rates. If true, this would further weaken the ability of reaching-for-yield to explain why the excess sensitivity of long rates has become so horizon-dependent.

Evidence from sectoral bond market flows. We use quarterly data from the Federal Reserve’s Financial Accounts on the aggregate net bond acquisitions by insurers, pension funds, and banks to construct proxies for the bond demand of yield-seeking investors, h_t . We focus on these intermediaries since prior research argues that they are likely to be concerned about the current yield on their portfolios and, thus, to reach for yield when interest rates fall.¹⁴ For intermediaries in sector i , we compute bond flows in quarter t as $\%FLOW_{i,t} = FLOW_{i,t}/HOLD_{i,t-1}$, where $FLOW_{i,t}$ denotes net bond acquisitions in quarter t and $HOLD_{i,t-1}$ is prior bond holdings. Bonds include the sum of U.S. Treasuries, agency debt and GSE-guaranteed mortgage-backed securities, and corporate bonds.

In the Internet Appendix, we estimate quarterly regressions that are analogous to equations (23) and (24) using these bond flows $\%FLOW_{i,t}$ as X_t . In the post-2000 data, we find little evidence that increases in short rates lead to reductions in bond purchases by insurers, pensions, and banks. Furthermore, bond purchases by these intermediaries do not predict low excess returns on long-term bonds.

In summary, we find evidence that mortgage refinancing and investor extrapolation both help explain why long yields rates have temporarily overreacted to short rates since 2000. However, we find little evidence that reaching-for-yield plays a major role in driving the temporary overreaction of long rates.¹⁵

¹⁴Insurers and banks are generally not required to include any changes in mark-to-market value in their reported earnings, potentially leading to yield-seeking behavior. For prior work on reaching-for-yield by insurers, see [Becker and Ivashina \(2015\)](#). For banks, see [Maddaloni and Peydró \(2011\)](#) and [Hanson and Stein \(2015\)](#). For pensions, see [Lu et al. \(2019\)](#).

¹⁵This need not imply that reaching-for-yield plays an unimportant role in determining financial risk premia more generally.

VI. IMPLICATIONS

VI.A. High-frequency identification

Our findings have clear implications for identification approaches based on the high-frequency responses of long-term yields to macroeconomic news and policy announcements. Papers in the vast event-study literature often implicitly assume that one can directly infer the expected long-run effects of news shocks on future fundamentals by looking at the high-frequency reactions of long-term asset prices (MacKinlay, 1997). And, several recent papers—e.g., Hördahl et al. (2015) and Nakamura and Steinsson (2018)—have made this assumption more explicitly. Intuitively, if changes in long rates in a tight window around a macro announcement are governed by the expectations hypothesis—e.g., because bond risk premia only vary at lower frequencies, then the high-frequency reaction of long rates directly reveals the expected long-run effect of the news shock on future short rates. For instance, if the 10-year forward rate fell by 20 basis points in a short window around an FOMC announcement, one would infer that this led expected short rates in 10 years to drop by 20 basis points.

Our evidence casts serious doubt on this assumption. If, as we argue, a large portion of the impact of news shock on long rates reflects rapidly-reverting shifts in term premia, then the short-run impact of news shocks on long rates will differ meaningfully from their expected long-run impact on future short rates. As a result, the high-frequency responses of long rates are likely to provide a highly biased estimate of the longer-run impact of announcements. Fortunately, it is relatively straightforward to eliminate this bias: one needs to use a methodology that does not assume that we can directly infer the expected long-run effects of news shocks simply by looking the high-frequency reactions of long-term asset prices.¹⁶ Of course, these unbiased approaches lead to far less precise estimates, so economists face a steep bias-variance trade-off. The short-run market impact of news on long rates can be estimated very precisely, but these are likely to be biased estimates of the longer-run impact that is typically of greatest interest.

Still, it is possible that changes in 1-year yields that coincide with macro announcements are different, and do not trigger movements in term premia, as argued by Hördahl et al. (2015) and Nakamura and Steinsson (2018). To provide some direct evidence, we form a macro news index for month t , $News_t$, by cumulating daily changes in 1-year yields within month t on days with important announcements. Our data on announcement timing is from Money Market Services/Action Economics and begins in 1980. The announcements we consider are: FOMC announcements, the

¹⁶For instance, one could estimate the long-run effects of a shock using a Structural VAR in which high-frequency asset prices movements are used as external instruments for monetary policy or other shocks. And, then IV-estimates of the SVAR would be used to trade out the long-run dynamic effects of the shock—see, e.g., Gertler and Karadi (2015) and Eberly et al. (2020). Similarly, one could estimate the long-run effects of news shocks using Jorda (2005) style “local projections” in which one regresses outcomes at future horizons on high-frequency market reactions to news.

employment situation report, retail sales, durable goods orders, new and existing home sales, housing starts, CPI, and PPI. We then estimate the following predictive regression for the subsequent change in 10-year forward rates:

$$(27) \quad f_{t+h}^{(10)} - f_t^{(10)} = \delta_0 + \delta_1 L_t + \delta_2 S_t + \delta_3 (L_t - L_{t-1}) + \delta_4 (S_t - S_{t-1}) + \lambda \cdot News_t + \varepsilon_{t+h},$$

where L_t and S_t denote the level and slope of the yield curve at the end of month t . Thus, equation (27) adds $News_t$ to the [Jorda \(2005\)](#) projections we previously estimated in equation (5). [Table VII](#) shows the results for both pre- and post-2000 samples and for $h = 3$ -, 6 -, 9 -, and 12 - month future changes.

In Panel A, we omit $News_t$, so the estimates are the same as those in [Table III](#). As previously shown, past increases in short-term rates are associated with predictable future declines in long-term forward rates in the post-2000 data, but there is no such tendency in the pre-2000 data. In Panel B, we add $News_t$, but omit the prior changes in level and slope. We see that positive values of the news index predict subsequent declines in long-term forwards in the post-2000 data. Indeed, the coefficients on $News_t$ in Panel B are similar to those on $L_t - L_{t-1}$ in Panel A.

In Panel C, we include $News_t$ as well as the prior changes in level and slope. The goal is to see if shifts in short rates on announcement and non-announcement days have different implications. Once we control for the total change in short rates in month t , $L_t - L_{t-1}$, we find that the coefficient on $News_t$ is small and insignificant, indicating that the response of long-term forwards rates on announcement days is just as likely to reverse as the response on non-announcement days.

In Panel D, we break $News_t$ into two pieces—one reflecting changes in short rates on FOMC announcement days and one for all other announcements—to see if FOMC announcements differ from other macro announcements. We exclude the 1980-81 monetary targeting regime and thus FOMC announcement dates begin in 1982. As in Panel C, we include $L_t - L_{t-1}$ as an independent variable. If anything, Panel D suggests that, since 2000, changes in short rates on FOMC announcement days are *more* likely to be followed by reversals in long-term forward rates than changes on non-announcement days.

VI.B. Monetary policy transmission

Our results also have important implications for the transmission of monetary policy. Central banks conduct conventional monetary policy by adjusting short-term nominal rates. According to the standard New Keynesian view ([Gali, 2008](#)), changes in nominal short rates affect real short rates because of nominal rigidities. And, the resulting shifts in real short rates affects long-term real rates via the expectations hypothesis, which then influence household consumption and firm investment. [Stein \(2013\)](#) points out that the excess sensitivity of long-term yields—whereby shocks to short rates move term premia in the same direction—should strengthen the effects of monetary

policy relative to the canonical view. [Stein \(2013\)](#) refers to this as the “recruitment” channel of monetary transmission.

In our framework from Section [IV.](#), the strength of this recruitment channel at business-cycle frequencies (e.g., over a 1 to 3-year horizon) depends on (i) the strength of the demand-based amplification mechanisms (i.e., the size of C) and (ii) the persistence of the associated demand shocks. When ρ_s is well below ρ_T as under the mortgage refinancing interpretation of C , the associated shifts in term premia would be quite transient and would likely have only modest effects on investment and consumption at medium-run frequencies. (A caveat here is that reductions in short rates that trigger mortgage refinancing waves may only temporarily lower term premia, but the effect of refinancing waves on distribution of household disposable income, and hence consumption, may persist long after term premia have reverted in heterogeneous agent settings.¹⁷) By contrast, when $\rho_s \approx \rho_T$ as under the reaching-for-yield interpretation, the shifts in term premia would be more persistent and have larger effects on aggregate demand.

Our results indicate that a significant part of the influence of short rates on term premia is quite transitory, suggesting that recruitment channel may be smaller than one would conclude based on a simple extrapolation of the high-frequency response of term premia to policy shocks documented by [Hanson and Stein \(2015\)](#), [Gertler and Karadi \(2015\)](#), and [Gilchrist et al. \(2015\)](#). More generally, our findings suggest that central banks should heed the way that monetary policy impacts financial conditions at business-cycle frequencies, but should focus less on the immediate market response to their announcements since much of the latter may be quite transitory. In this way, our findings lend support to the argument in [Stein and Sunderam \(2018\)](#) that the Fed has become too focused on high-frequency asset price movements.

VI.C. Bond market “conundrums”

Our findings help explain the rising prevalence of episodes like the one [Greenspan \(2005\)](#) famously called the “conundrum”—the period after June 2004 when short rates rose and long rates fell. Consistent with the weaker low-frequency sensitivity of long rates, “conundrum” episodes—defined as 6-month periods where short and long rates move in *opposite* directions—have grown increasingly common. Since 2000, 1- and 10-year yields have moved in the *opposite* direction in 37% of all 6-month periods. By contrast, from 1971 to 1999, this figure was 18%, and the difference is significant ($p\text{-val} < 0.001$). In the Internet Appendix, we show that the non-Markovian dynamics documented in Section [III.](#) help explain several noteworthy “conundrums” episodes, including Greenspan’s 2004 “conundrum,” 2008, and 2017. In each case, 1-year and 10-year yields moved in opposite directions, but, if the slope of the yield curve had not responded to past changes short

¹⁷To the extent that mortgage refinancing plays an important role in U.S. monetary policy transmission as in recent heterogeneous agent models ([Beraja et al., 2018](#); [Berger et al., 2018](#); [Wong, 2019](#)), then even short-lived excess sensitivity may make monetary policy more potent than in a world where long rates are not excessively sensitive.

rates, 10- and 1-year yields would have moved in the same direction.

VI.D. Affine term structure models

Finally, we explore the implications for affine term structure models which are a widely-used, reduced-form tools for understanding the term structure of bond yields (Duffie and Kan, 1996; Duffee, 2002). In these models, the n -year zero coupon yield is $y_t^{(n)} = \alpha_{0(n)} + \alpha'_{1(n)}\mathbf{x}_t$, where \mathbf{x}_t is a vector of state variables and the $\alpha_{0(n)}$ and $\alpha_{1(n)}$ coefficients satisfy a set of recursive equations. In the Internet Appendix, we fit affine term structure models using the first K principal components of 1- to 10-year yields as the state variables \mathbf{x}_t . We show that standard affine models—models that are Markovian with respect to these current yield-curve factors—cannot fit the fact that β_h declines so strongly with horizon h in the post-2000 data. This remains so even if we estimate models that include many (e.g., $K = 5$) yield-curve factors as state variables. However, we show that our key finding is consistent with non-Markovian term structure models in which past lags of the yield-curve factors are treated as “unspanned state variables.”¹⁸

VII. CONCLUSION

The strong sensitivity of long-term interest rates to changes in short rates is a long-standing puzzle. We have shown that since 2000 this sensitivity has become even stronger at high frequencies, but has weakened significantly at low-frequencies. As a result, in the post-2000 data, the sensitivity of long rates to changes in short rates declines steeply with the horizon over which these changes are computed.

Before 2000, long rates were quite sensitive to short rates because inflation expectations were relatively unanchored, making short rates highly persistent. Since 2000, the sensitivity of long rates has become horizon-dependent and arises because past increases in short rates temporarily raise the term premium, leading long rates to temporarily overreact to changes in short rates. Consistent with this view, we show that, controlling for current yields, past changes in short rates predict future yield-curve flattening, declines in long-term yields and forwards, and high excess returns on long-term bonds after 2000.

We proposed a model that can explain this puzzling post-2000 pattern. The tendency of long rates to temporarily overreact to changes in short rates is due to the combination of (i) rate-amplifying shifts in the demand for long-term bonds and (ii) a limited and slow arbitrage response to these demand shifts. We presented evidence that two specific rate-amplifying demand mechanisms—mortgage refinancing waves and extrapolation of past changes in short rates—each

¹⁸An unspanned state variable is a variable that is useful for forecasting future bond yields and returns but that has no impact on the current yield curve (Duffee, 2002). To be clear, we do not argue that the past increase in the level of rates is *literally* unspanned. Instead, as discussed in the Internet Appendix, we think this variable is *close* to being unspanned.

help explain this post-2000 pattern.

Our findings have important implications for the recruitment channel of monetary policy transmission (Stein, 2013). In recent years this channel appears far more short-lived than one might conclude from high-frequency evidence alone: Part of the high-frequency response of long rates to shocks to short rates represents transitory term premium movements. Lastly, it is important to remember that event-study approaches only measure high-frequency responses of long-term rates to news; the impact may be more muted at the lower frequencies that are typically of greatest interest to economists and policymakers.

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TABLE I
Regressions of changes in long-term rates on short-term

PANEL A. 10-year zero coupon yields and IC				
	(1)	(2)	(3)	(4)
	Nominal	Nominal	Real	IC
Daily	0.56*** [0.02]	0.87*** [0.03]	0.54*** [0.03]	0.33*** [0.02]
Monthly	0.46*** [0.04]	0.66*** [0.11]	0.38*** [0.09]	0.26*** [0.09]
Quarterly	0.48*** [0.04]	0.44*** [0.07]	0.22** [0.10]	0.22* [0.12]
Semi-annual	0.50*** [0.04]	0.34*** [0.07]	0.21** [0.08]	0.13 [0.09]
Yearly	0.56*** [0.05]	0.23*** [0.05]	0.15** [0.06]	0.08 [0.05]
Sample	1971-1999	2000-2019	2000-2019	2000-2019
PANEL B. 10-year instantaneous forward yields and IC				
	(1)	(2)	(3)	(4)
	Nominal	Nominal	Real	IC
Daily	0.39*** [0.03]	0.49*** [0.04]	0.31*** [0.03]	0.17*** [0.03]
Monthly	0.29*** [0.04]	0.26* [0.14]	0.18** [0.08]	0.06 [0.09]
Quarterly	0.31*** [0.05]	0.06 [0.09]	0.09* [0.05]	-0.03 [0.05]
Semi-annual	0.33*** [0.06]	-0.02 [0.08]	0.04 [0.04]	-0.06 [0.05]
Yearly	0.39*** [0.07]	-0.13** [0.06]	-0.02 [0.04]	-0.11** [0.04]
Sample	1971-1999	2000-2019	2000-2018	2000-2019

Notes: This table reports the estimated regression coefficients from equations (1) and (2) for each reported sample. The dependent variable is the change in the 10-year U.S. Treasury zero-coupon yield or forward rate, either nominal, real or their difference (IC, or inflation compensation). The independent variable is the change in the 1-year nominal U.S. Treasury zero-coupon yield in all cases. Changes are considered with daily data, and with monthly data using monthly ($h = 1$), quarterly ($h = 3$), semi-annual ($h = 6$) and annual ($h = 12$) horizons. In the 1971-1999 monthly sample, time t runs from Aug-1971 to Dec-1999 and the number of monthly observations is 341 irrespective of h . In the 2000-2019 monthly sample, t runs from Jan-2000 to Dec-2019, so the number of monthly observations runs 239 from for $h = 1$ to 228 for $h = 12$. For $h > 1$, we report Newey-West (1987) standard errors are in brackets, using a lag truncation parameter of $\lceil 1.5 \times h \rceil$; for $h = 1$, we report heteroskedasticity robust standard errors. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Significance is computed using the asymptotic theory of [Kiefer and Vogelsang \(2005\)](#) which has better finite sample properties than traditional asymptotic theory.

TABLE II
Estimates of predictive equations for level and slope

	Pre-2000			Post-2000		
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable: Level						
L_t	0.98*** [0.02]	0.97*** [0.02]	0.96*** [0.02]	0.97*** [0.01]	0.98*** [0.01]	0.98*** [0.01]
S_t	0.00 [0.04]	-0.01 [0.04]	-0.02 [0.04]	-0.02* [0.01]	-0.01 [0.01]	-0.00 [0.01]
$L_t - L_{t-6}$		-0.01 [0.04]	0.05 [0.05]		0.08*** [0.03]	0.06** [0.03]
$S_t - S_{t-6}$			0.13* [0.07]			-0.03* [0.02]
Dependent Variable: Slope						
L_t	0.01 [0.01]	0.01 [0.01]	0.01 [0.01]	0.00 [0.01]	-0.01 [0.01]	-0.01 [0.01]
S_t	0.96*** [0.03]	0.96*** [0.02]	0.97*** [0.03]	0.98*** [0.02]	0.96*** [0.02]	0.96*** [0.02]
$L_t - L_{t-6}$		0.00 [0.02]	-0.03 [0.03]		-0.11*** [0.02]	-0.12*** [0.03]
$S_t - S_{t-6}$			-0.08 [0.05]			-0.02 [0.03]
N	341	335	335	239	239	239
Implied β_1	0.46	0.46	0.46	0.66	0.71	0.71
Implied β_{12}	0.52	0.51	0.58	0.59	0.38	0.30
AIC	-5720.3	-5607.9	-5608.7	-4529.0	-4567.0	-4565.6
BIC	-5697.3	-5577.4	-5570.5	-4508.1	-4539.2	-4530.8
Sample	1971-1999	1972-1999	1972-1999	2000-2019	2000-2019	2000-2019

Notes: This table reports the estimated regression coefficients from monthly predictive equations (3a) and (3b) for the Aug-1971 to Dec-1999 and Jan-2000 to Dec-2019 samples. Dependent variables are the level ($L_t \equiv y_t^{(1)}$) and slope ($S_t \equiv y_t^{(10)} - y_t^{(1)}$) of the U.S. Treasury zero-coupon yield curve. Heteroskedasticity robust standard errors are in brackets. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The table also shows AIC and BIC values (to be minimized) for each possible specification of the system of two equations. Lastly, the implied β_1 and β_{12} coefficients from equation (4) for each possible specification of the system are reported.

TABLE III
Predictable yield-curve dynamics following an impulse to short-term interest rates

	Pre-2000				Post-2000			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dep. var with h	3	6	9	12	3	6	9	12
Dependent variable: $y_{t+h}^{(10)} - y_t^{(10)}$								
$L_t - L_{t-1}$	-0.03 [0.12]	0.10 [0.13]	0.04 [0.11]	0.23* [0.12]	-0.11 [0.13]	-0.36* [0.19]	-0.41* [0.20]	-0.20 [0.20]
$S_t - S_{t-1}$	0.03 [0.18]	0.30 [0.19]	0.34 [0.20]	0.55** [0.26]	-0.09 [0.15]	-0.16 [0.17]	-0.11 [0.15]	-0.16 [0.20]
L_t	-0.06* [0.03]	-0.11* [0.06]	-0.17* [0.08]	-0.23** [0.10]	-0.07** [0.03]	-0.14*** [0.05]	-0.19*** [0.06]	-0.23*** [0.07]
S_t	-0.13* [0.07]	-0.27** [0.11]	-0.45*** [0.16]	-0.62*** [0.19]	-0.10** [0.05]	-0.20** [0.09]	-0.28** [0.11]	-0.35** [0.12]
Adj. R^2	0.03	0.08	0.14	0.20	0.05	0.11	0.16	0.20
Dependent variable: $f_{t+h}^{(10)} - f_t^{(10)}$								
$L_t - L_{t-1}$	-0.08 [0.12]	0.09 [0.11]	-0.05 [0.10]	0.12 [0.12]	-0.30*** [0.11]	-0.52*** [0.17]	-0.71*** [0.14]	-0.74*** [0.15]
$S_t - S_{t-1}$	0.01 [0.18]	0.18 [0.19]	0.16 [0.20]	0.34 [0.22]	-0.14 [0.16]	-0.11 [0.20]	-0.02 [0.19]	-0.17 [0.21]
L_t	-0.04 [0.03]	-0.09 [0.05]	-0.14* [0.08]	-0.19** [0.09]	-0.03 [0.03]	-0.05 [0.05]	-0.06 [0.06]	-0.06 [0.07]
S_t	-0.16** [0.06]	-0.31*** [0.11]	-0.48*** [0.15]	-0.66*** [0.19]	-0.09* [0.05]	-0.17* [0.09]	-0.27** [0.11]	-0.35** [0.14]
Adj. R^2	0.04	0.10	0.17	0.24	0.04	0.09	0.17	0.23
Dependent variable: $L_{t+h} - L_t$								
$L_t - L_{t-1}$	0.07 [0.24]	0.26 [0.21]	0.18 [0.20]	0.60*** [0.19]	0.66*** [0.17]	0.86** [0.38]	1.33** [0.53]	1.70*** [0.59]
$S_t - S_{t-1}$	0.13 [0.30]	0.64** [0.30]	0.57* [0.31]	1.07** [0.49]	-0.23** [0.11]	-0.38* [0.21]	-0.59* [0.29]	-0.58 [0.39]
L_t	-0.09* [0.05]	-0.18** [0.08]	-0.26** [0.11]	-0.36*** [0.12]	-0.08** [0.03]	-0.19** [0.07]	-0.28*** [0.11]	-0.38** [0.15]
S_t	-0.00 [0.11]	-0.09 [0.15]	-0.24 [0.20]	-0.33 [0.25]	-0.03 [0.04]	-0.07 [0.09]	-0.04 [0.15]	-0.01 [0.21]
Adj. R^2	0.03	0.08	0.12	0.17	0.20	0.26	0.36	0.42
Dependent variable: $S_{t+h} - S_t$								
$L_t - L_{t-1}$	-0.10 [0.16]	-0.16 [0.12]	-0.14 [0.12]	-0.37*** [0.11]	-0.77*** [0.16]	-1.23*** [0.33]	-1.74*** [0.39]	-1.90*** [0.51]
$S_t - S_{t-1}$	-0.10 [0.17]	-0.34* [0.18]	-0.23 [0.16]	-0.51 [0.30]	0.14 [0.12]	0.22 [0.22]	0.48* [0.27]	0.43 [0.32]
L_t	0.03 [0.03]	0.07* [0.03]	0.10** [0.04]	0.13*** [0.04]	0.01 [0.03]	0.05 [0.07]	0.09 [0.10]	0.15 [0.14]
S_t	-0.12* [0.07]	-0.18* [0.10]	-0.21* [0.11]	-0.29** [0.13]	-0.07* [0.04]	-0.13 [0.09]	-0.24 [0.13]	-0.34 [0.20]
Adj. R^2	0.08	0.16	0.20	0.28	0.16	0.24	0.37	0.43
N	340	340	340	340	237	234	231	228
Sample	1971-1999	1971-1999	1971-1999	1971-1999	2000-2019	2000-2019	2000-2019	2000-2019

Notes: This table reports the estimated regression coefficients in equation (5) for the Aug-1971 to Dec-1999 and Jan-2000 to Dec-2019 samples. For $h = 3, 6, 9,$ and, 12-months changes, we show results for 10-year yields ($z_t = y_t^{(10)}$), 10-year forward rates ($z_t = f_t^{(10)}$), level ($z_t = L_t$), and slope ($z_t = S_t$). We report Newey-West standard errors in brackets using a lag truncation parameter of $\lceil 1.5 \times h \rceil$. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Significance is computed using the asymptotic theory of [Kiefer and Vogelsang \(2005\)](#) which has better finite sample properties than traditional asymptotic theory.

TABLE IV
Estimates of predictive equations for bond excess returns

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$rx_{t \rightarrow t+k}^{(10)}$ for $k =$	1	1	1	3	3	3	6	6	6
Panel A: Pre-2000									
L_t	0.17 [0.13]	0.18 [0.13]	0.22* [0.12]	0.53* [0.31]	0.54 [0.33]	0.65* [0.34]	0.99* [0.57]	1.01 [0.62]	1.25** [0.60]
S_t	0.55** [0.24]	0.61** [0.24]	0.66*** [0.24]	1.57** [0.64]	1.91*** [0.65]	2.07*** [0.64]	3.17*** [1.04]	3.62*** [1.13]	3.98*** [1.09]
$L_t - L_{t-6}$		0.10 [0.19]	-0.18 [0.29]		0.59 [0.49]	-0.22 [0.54]		0.79 [0.69]	-1.00 [0.90]
$S_t - S_{t-6}$			-0.57 [0.44]			-1.67** [0.81]			-3.67*** [1.30]
Adj. R^2	0.02	0.01	0.02	0.05	0.06	0.07	0.11	0.12	0.17
N	341	335	335	341	335	335	341	335	335
Sample	1971-1999	1972-1999	1972-1999	1971-1999	1972-1999	1972-1999	1971-1999	1972-1999	1972-1999
Panel B: Post-2000									
L_t	0.30*** [0.11]	0.32*** [0.11]	0.28** [0.11]	0.73*** [0.27]	0.81*** [0.28]	0.71** [0.30]	1.37*** [0.44]	1.50*** [0.47]	1.34** [0.51]
S_t	0.56*** [0.21]	0.63*** [0.20]	0.53** [0.21]	1.41*** [0.48]	1.65*** [0.52]	1.42** [0.58]	2.71*** [0.83]	3.07*** [0.93]	2.73*** [0.95]
$L_t - L_{t-6}$		0.33 [0.24]	0.62* [0.33]		0.98** [0.46]	1.66*** [0.61]		1.33* [0.73]	2.39** [1.07]
$S_t - S_{t-6}$			0.48 [0.35]			1.12 [0.72]			1.76 [1.22]
Adj. R^2	0.03	0.03	0.03	0.08	0.10	0.12	0.16	0.18	0.20
N	239	239	239	237	237	237	234	234	234
Sample	2000-2019	2000-2019	2000-2019	2000-2019	2000-2019	2000-2019	2000-2019	2000-2019	2000-2019

Notes: This table reports the estimated regression coefficients in equation (7) using monthly data from the Aug-1971 to Dec-1999 and Jan-2000 to Dec-2019 samples. We report results various return forecast horizon (k). The yield on k -month Treasury bills, $y_t^{(k/12)}$, is from the yield curve estimates in [Gürkaynak et al. \(2007\)](#). However, this curve is based on coupon securities with at least three months to maturity and does not fit the very short end of the curve well in the pre-2000 data. Therefore, we take the 1-month bill yield from Ken French's website for the pre-2000 sample. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. For $k = 1$ -month returns, we report heteroskedasticity robust standard errors are in brackets. For $k = 3$ and 6-month returns, we report [Newey and West \(1987\)](#) standard errors in brackets, using a lag truncation parameter of 5 and 9 months, respectively. In this case, p -values are computed using the asymptotic theory of [Kiefer and Vogelsang \(2005\)](#) which has better finite sample properties than traditional asymptotic theory.

TABLE V
The role of mortgage refinancing: Evidence from mortgage-related quantities

	(1)	(2)	(3)	(4)	(5)	(6)	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Dependent Variable: $y_t^M - c_t^M$						Panel B. Dependent Variable: DU_t^M						
L_t	0.76*** [0.07]	0.76*** [0.06]	0.72*** [0.06]	0.23*** [0.05]	0.24*** [0.05]	0.20*** [0.05]	0.16*** [0.04]	0.16*** [0.04]	0.15*** [0.04]	0.05 [0.09]	0.08 [0.09]	0.01 [0.08]
S_t	0.17 [0.10]	0.13 [0.11]	0.05 [0.10]	-0.03 [0.07]	0.00 [0.06]	-0.08 [0.06]	0.15** [0.08]	0.08 [0.08]	0.06 [0.08]	-0.06 [0.14]	0.03 [0.12]	-0.11 [0.12]
$L_t - L_{t-6}$		-0.07 [0.08]	0.18* [0.10]	0.17** [0.08]	0.42*** [0.07]	0.42*** [0.06]	-0.13** [0.05]	-0.06 [0.06]	0.17** [0.08]	0.44*** [0.13]	0.90*** [0.12]	0.90*** [0.12]
$S_t - S_{t-6}$			0.56*** [0.16]									0.76*** [0.12]
Adj. R^2	0.86	0.86	0.88	0.52	0.55	0.67	0.23	0.29	0.30	0.02	0.16	0.36
Panel C. Dependent Variable: $rx_{t \rightarrow t+3}^{(10)}$						Panel D. Dependent Variable: $rx_{t \rightarrow t+3}^{(10)}$						
L_t	0.67* [0.36]	1.10* [0.59]	0.86 [0.59]	0.71** [0.30]	0.21 [0.30]	0.34 [0.34]	0.67* [0.36]	0.82** [0.36]	0.88** [0.36]	0.71** [0.30]	0.68** [0.28]	0.70** [0.30]
S_t	2.20** [0.75]	1.83** [0.77]	2.21*** [0.76]	1.42** [0.58]	1.50*** [0.49]	1.57*** [0.57]	2.20*** [0.75]	1.98*** [0.71]	2.28*** [0.73]	1.42** [0.58]	1.51*** [0.50]	1.55*** [0.58]
$L_t - L_{t-6}$	-0.30 [0.56]	-0.25 [0.60]	-0.25 [0.60]	1.66*** [0.61]	0.88 [0.79]	0.88 [0.79]	-0.30 [0.56]	-0.38 [0.55]	-0.38 [0.55]	1.66*** [0.61]	0.71 [0.83]	0.71 [0.83]
$S_t - S_{t-6}$	-1.70* [0.92]	-1.55 [1.08]	-1.55 [1.08]	1.12 [0.72]	0.35 [0.80]	0.35 [0.80]	-1.70* [0.92]	-1.45 [0.90]	-1.45 [0.90]	1.12 [0.72]	0.31 [0.85]	0.31 [0.85]
$y_t^M - c_t^M$		-0.74 [0.71]	-0.27 [0.77]		2.34*** [0.80]	1.82* [0.98]		-1.77* [0.94]	-1.45* [0.81]		1.31*** [0.35]	1.05** [0.51]
Adj. R^2	0.07	0.05	0.06	0.12	0.13	0.14	0.07	0.07	0.08	0.12	0.14	0.14
Sample	1976-1999	1976-1999	1976-1999	2000-2019	2000-2019	2000-2019	1976-1999	1976-1999	1976-1999	2000-2019	2000-2019	2000-2019

Notes: Panels A and B report the estimated coefficients for equation (23) where the dependent variable is the mortgage refinancing disincentive: $X_t = y_t^M - c_t^M$ or the duration of the Barclay's MBS index: $X_t = DU_t^M$. Panels C and D report the estimated coefficients using equation (24) to forecast 3-month returns using either the mortgage refinancing disincentive or the duration. All regressions are estimated using monthly data for the Jan-1976 to Dec-1999 and Jan-2000 to Dec-2019 samples. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. We report Newey and West (1987) standard errors in brackets, using a lag truncation parameter of 9 months in Panels A and B and 5 months in Panels C and D. p -values are computed using the asymptotic theory of Kiefer and Vogelsang (2005) which has better finite sample properties than traditional asymptotic theory.

TABLE VI
The role of investor over-extrapolation: Evidence from bond mutual fund flows

	Pre-2000			Post-2000		
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable: $FLOW_{t-3 \rightarrow t}$						
L_t	1.96** [0.89]	2.15*** [0.67]	2.24*** [0.69]	0.06 [0.23]	0.01 [0.20]	0.04 [0.23]
S_t	4.08** [1.64]	3.24** [1.27]	3.92** [1.45]	0.60 [0.37]	0.44 [0.36]	0.51 [0.39]
$L_t - L_{t-6}$		-4.00*** [1.04]	-5.29*** [1.29]		-0.77** [0.32]	-1.00** [0.46]
$S_t - S_{t-6}$			-4.31** [2.15]			-0.38 [0.45]
Adj. R^2	0.24	0.43	0.50	0.09	0.16	0.16
N	189	189	189	240	240	240
Dependent Variable: $rx_{t \rightarrow t+3}^{(10)}$						
L_t	1.27*** [0.30]	1.13*** [0.29]	1.04*** [0.35]	0.71** [0.30]	0.76*** [0.27]	0.72** [0.31]
S_t	1.88*** [0.66]	1.37** [0.64]	1.47* [0.74]	1.42** [0.58]	1.63*** [0.50]	1.55** [0.60]
$L_t - L_{t-6}$	-0.01 [0.61]		0.54 [0.89]	1.66*** [0.61]		1.42** [0.64]
$S_t - S_{t-6}$	-1.01 [0.88]		-0.56 [0.98]	1.12 [0.72]		1.02 [0.73]
$FLOW_{t-3 \rightarrow t}$		0.07 [0.08]	0.10 [0.11]		-0.34** [0.14]	-0.24 [0.15]
Adj. R^2	0.16	0.17	0.17	0.12	0.10	0.12
N	189	189	189	237	237	237
Sample	1984-1999	1984-1999	1984-1999	2000-2019	2000-2019	2000-2019

Notes: Data on flows into taxable bond mutual funds is from the Investment Company Institute. Letting $\%FLOW_t = FLOW_t/TNA_{t-1}$ denote the percentage flow in month t , the 3-month cumulative percentage flow is $\%FLOW_{t-3 \rightarrow t} = (1 + \%FLOW_t)(1 + \%FLOW_{t-1})(1 + \%FLOW_{t-2}) - 1$. Panels A reports the estimated regression coefficients for equation (23) using $X_t = \% \Delta FLOW_{t-3 \rightarrow t}$. Panels B and C report the estimated regression coefficients when we use $X_t = \% \Delta FLOW_{t-3 \rightarrow t}$ in equation (24) to forecast 3-month returns. We estimate these regressions using monthly data for the Apr1984 to Dec-1999 and Jan-2000 to Dec-2019 subsamples. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. We report Newey and West (1987) standard errors in brackets, using a lag truncation parameter of 9 months in Panel A and 5 months in Panel B. p -values are computed using the asymptotic theory of Kiefer and Vogelsang (2005) which has better finite sample properties than traditional asymptotic theory.

TABLE VII
Economic news and subsequent changes in forward rates

	Pre-2000				Post-2000			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$f_{t+h}^{(10)} - f_t^{(10)}$ with h	3	6	9	12	3	6	9	12
Panel A								
$L_t - L_{t-1}$	-0.09 [0.12]	0.11 [0.12]	-0.04 [0.12]	0.12 [0.13]	-0.30*** [0.11]	-0.52*** [0.17]	-0.71*** [0.16]	-0.74*** [0.15]
$S_t - S_{t-1}$	0.03 [0.19]	0.29 [0.20]	0.19 [0.22]	0.35 [0.24]	-0.14 [0.16]	-0.11 [0.20]	-0.02 [0.20]	-0.17 [0.21]
L_t	-0.04 [0.03]	-0.08 [0.06]	-0.13 [0.08]	-0.18* [0.09]	-0.03 [0.03]	-0.05 [0.05]	-0.06 [0.06]	-0.06 [0.07]
S_t	-0.11 [0.07]	-0.24* [0.13]	-0.37** [0.17]	-0.52** [0.21]	-0.09* [0.05]	-0.17* [0.09]	-0.27** [0.11]	-0.35** [0.14]
Adj. R^2	0.02	0.05	0.09	0.14	0.04	0.09	0.17	0.23
Panel B								
$News_t$	-0.29 [0.21]	-0.12 [0.30]	-0.80*** [0.24]	-0.46 [0.26]	-0.45** [0.18]	-0.63* [0.34]	-0.82*** [0.27]	-0.67** [0.30]
L_t	-0.04 [0.03]	-0.08 [0.06]	-0.12 [0.08]	-0.17* [0.09]	-0.04 [0.03]	-0.05 [0.05]	-0.06 [0.06]	-0.06 [0.07]
S_t	-0.10 [0.07]	-0.23* [0.12]	-0.37* [0.17]	-0.52** [0.20]	-0.10** [0.04]	-0.18** [0.08]	-0.26** [0.11]	-0.35** [0.13]
Adj. R^2	0.02	0.05	0.11	0.14	0.04	0.08	0.15	0.21
Panel C								
$News_t$	-0.11 [0.34]	-0.11 [0.35]	-0.99** [0.42]	-0.72** [0.29]	-0.28 [0.29]	-0.05 [0.47]	0.15 [0.36]	0.38 [0.29]
$L_t - L_{t-1}$	-0.06 [0.16]	0.14 [0.15]	0.19 [0.16]	0.29 [0.17]	-0.19 [0.17]	-0.50* [0.24]	-0.77*** [0.22]	-0.89*** [0.18]
$S_t - S_{t-1}$	0.03 [0.19]	0.29 [0.20]	0.23 [0.20]	0.38 [0.23]	-0.15 [0.16]	-0.11 [0.20]	-0.02 [0.18]	-0.15 [0.21]
L_t	-0.04 [0.03]	-0.08 [0.06]	-0.12 [0.08]	-0.17* [0.09]	-0.03 [0.03]	-0.05 [0.05]	-0.06 [0.06]	-0.06 [0.07]
S_t	-0.11 [0.07]	-0.24* [0.13]	-0.37* [0.17]	-0.52** [0.21]	-0.09* [0.05]	-0.17* [0.09]	-0.27** [0.11]	-0.35** [0.14]
Adj. R^2	0.01	0.05	0.10	0.14	0.04	0.08	0.16	0.23
N	240	240	240	240	237	234	231	228
Sample	1980-1999	1980-1999	1980-1999	1980-1999	2000-2019	2000-2019	2000-2019	2000-2019
Panel D								
$News_{t,FOMC}$	0.11 [0.77]	-0.01 [1.04]	-1.89 [1.56]	-0.44 [1.56]	-0.99** [0.46]	-0.74* [0.39]	-1.18 [0.75]	0.09 [0.56]
$News_{t,Other}$	-0.25 [0.33]	-0.17 [0.45]	-0.45 [0.48]	-0.32 [0.43]	-0.02 [0.29]	0.21 [0.57]	0.65 [0.41]	0.49 [0.41]
$L_t - L_{t-1}$	0.09 [0.18]	0.03 [0.21]	0.03 [0.24]	-0.00 [0.27]	-0.20 [0.16]	-0.51** [0.24]	-0.79*** [0.20]	-0.89*** [0.18]
$S_t - S_{t-1}$	0.20 [0.24]	0.30 [0.27]	0.33 [0.23]	0.55** [0.25]	-0.17 [0.16]	-0.13 [0.20]	-0.05 [0.18]	-0.16 [0.21]
L_t	-0.07* [0.03]	-0.13* [0.06]	-0.18* [0.09]	-0.20* [0.10]	-0.03 [0.03]	-0.05 [0.05]	-0.06 [0.06]	-0.06 [0.07]
S_t	-0.07 [0.07]	-0.17 [0.14]	-0.28 [0.19]	-0.42* [0.21]	-0.10** [0.05]	-0.18* [0.09]	-0.28** [0.11]	-0.36** [0.14]
Adj. R^2	0.03	0.08	0.14	0.16	0.05	0.09	0.18	0.23
N	216	216	216	216	237	234	231	228
Sample	1982-1999	1982-1999	1982-1999	1982-1999	2000-2019	2000-2019	2000-2019	2000-2019

Notes: This table reports the regression coefficients in equation (27) using monthly data from the Aug1971 to Dec-1999 and Jan-2000 to Dec2019 samples. Newey-West (1987) standard errors are in brackets, using a lag truncation parameter of $\lceil 1.5 \times h \rceil$. Significance: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Significance is computed using the asymptotic theory of Kiefer and Vogelsang (2005) which has better finite sample properties than traditional asymptotic theory.

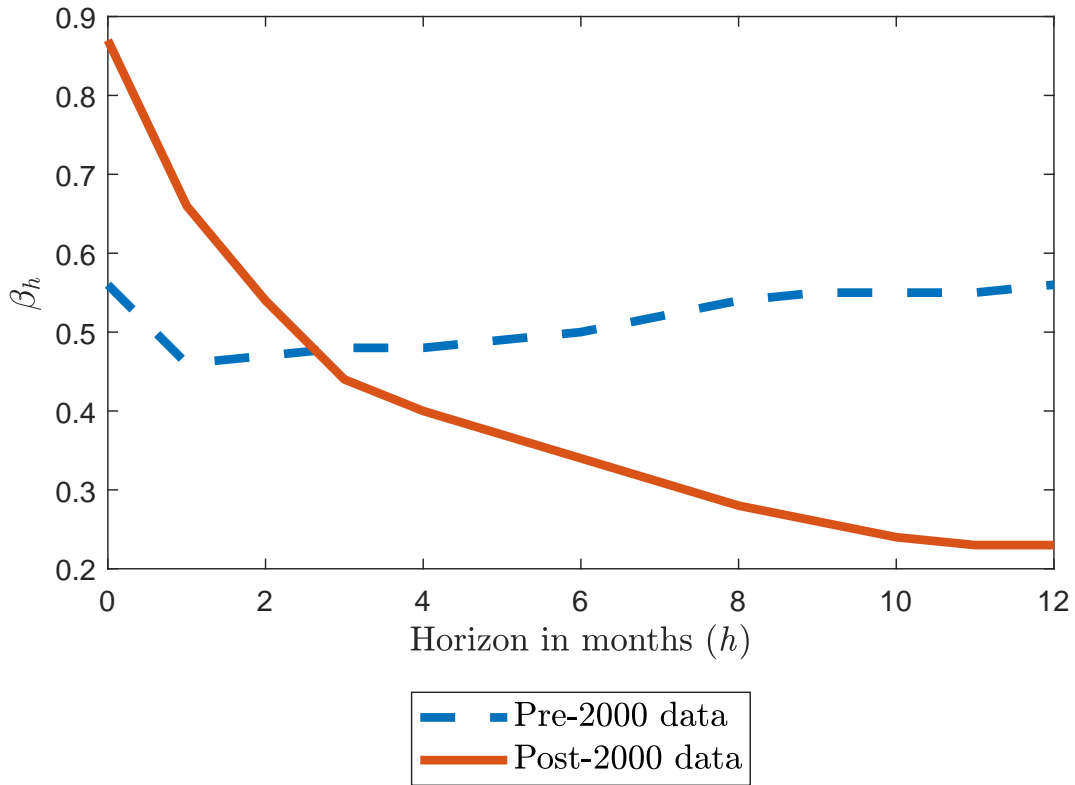


FIGURE I
Regressions of changes in long-term yields on short-term rates

This figure plots the estimated regression coefficients β_h from equation (1) versus horizon (h) for the pre-2000 and post-2000 sample: $y_{t+h}^{(10)} - y_t^{(10)} = \alpha_h + \beta_h(y_{t+h}^{(1)} - y_t^{(1)}) + \varepsilon_{t,t+h}$. The dependent variable is the h -month change in the 10-year nominal zero-coupon U.S. Treasury yield and the independent variable is the h -month change in the 1-year nominal zero-coupon U.S. Treasury yield. Changes are considered with daily data (plotted as $h = 0$ in the figure) and with monthly data using $h = 1, \dots, 12$ -month changes.

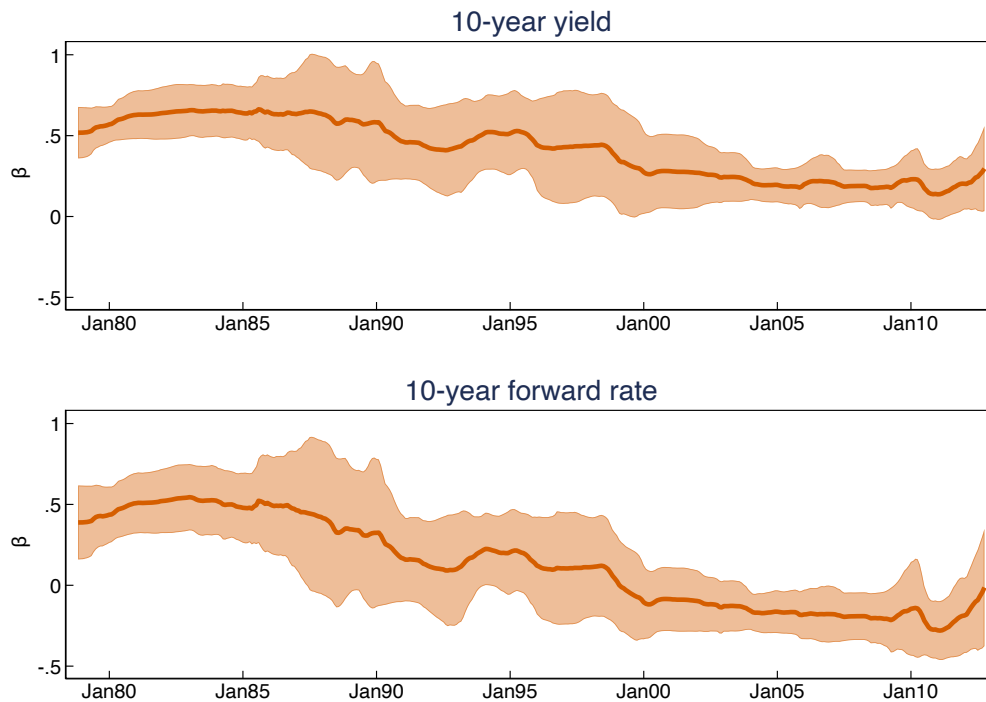


FIGURE II
Rolling regression estimates of equations (1) and (2)

This figure plots rolling estimates of the slope coefficients in equations (1) and (2) with $h = 12$ -month changes using 10-year rolling windows for estimation. Results are plotted against the midpoint of the 10-year rolling window. 95% confidence intervals are included (shaded areas), formed using Newey-West standard errors with a lag truncation parameter of 18 and 95% critical values from the asymptotic theory of [Kiefer and Vogelsang \(2005\)](#). Specifically, the 95% confidence interval is ± 2.41 times the estimated standard errors as opposed to ± 1.96 under traditional asymptotic theory.

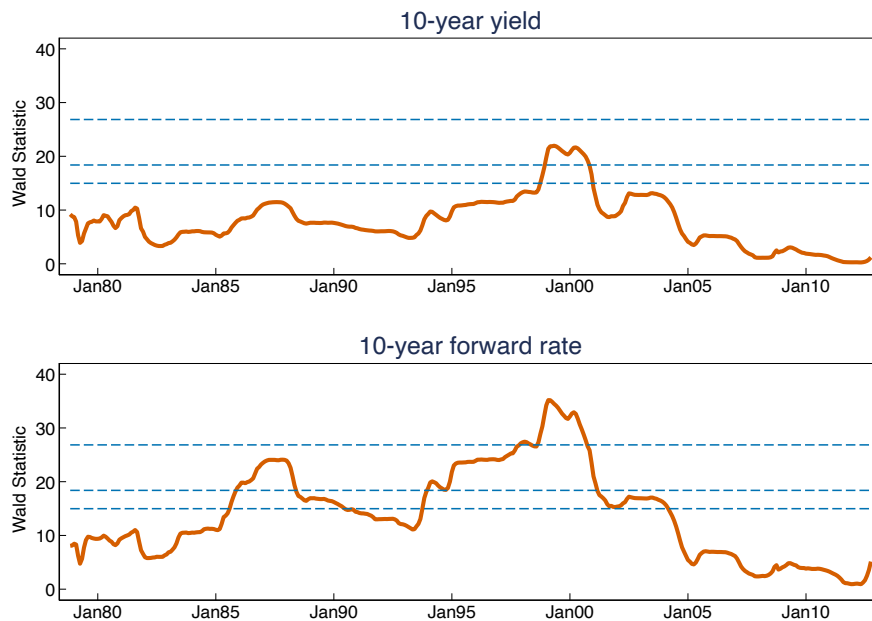


FIGURE III
Break tests for equations (1) and (2)

This figure plots the Wald test statistic for each possible break date in equations (1) and (2) with $h = 12$ -month changes from a fraction 15% of the way through the sample to 85% of the way through the sample. The horizontal red dashed lines denote 10%, 5%, and 1% critical values for the maximum of these Wald statistics as in Andrews (1993). Our Wald tests use a Newey and West (1987) variance matrix with a lag truncation parameter of 18. To address the tendency for tests based on the Newey-West variance estimator to over-reject in finite samples, we use the Cho and Vogelsang (2017) critical values for a null of no structural break. The Cho and Vogelsang (2017) critical values are based on the asymptotic theory of Kiefer and Vogelsang (2005) and are slightly larger than the traditional critical values from Andrews (1993).

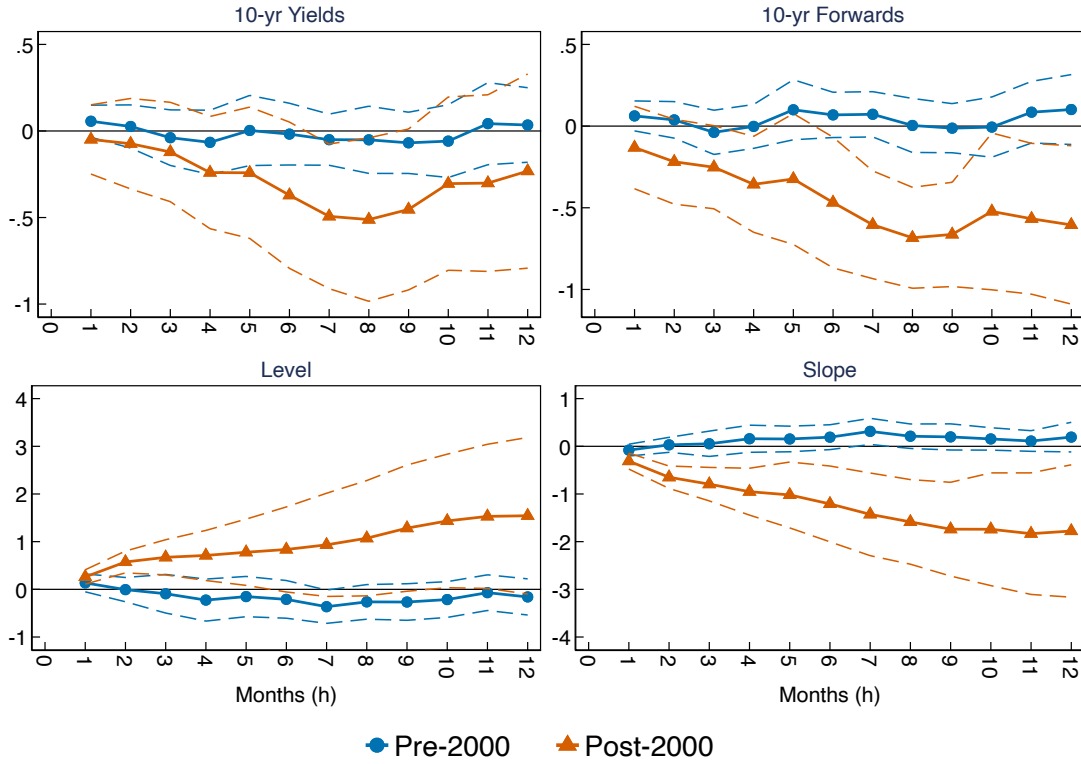


FIGURE IV

Predictable yield-curve dynamics following an impulse to short-term interest rates

The figures plot the coefficients $\delta_3^{(h)}$ versus horizon h from estimating equations (5) for various horizons $h = 1, \dots, 12$ -months in the pre-2000 and post-2000 samples. We show results for 10-year yields ($z_t = y_t^{(10)}$), 10-year forward rates ($z_t = f_t^{(10)}$), level ($z_t = L_t$) and slope ($z_t = S_t$). 95% confidence intervals are shown as dashed lines, formed using Newey-West standard errors and 95% critical values from the asymptotic theory of [Kiefer and Vogelsang \(2005\)](#). We use a Newey-West lag truncation parameter of 0 for $h = 1$ and $\lceil 1.5 \times h \rceil$ for $h > 1$.

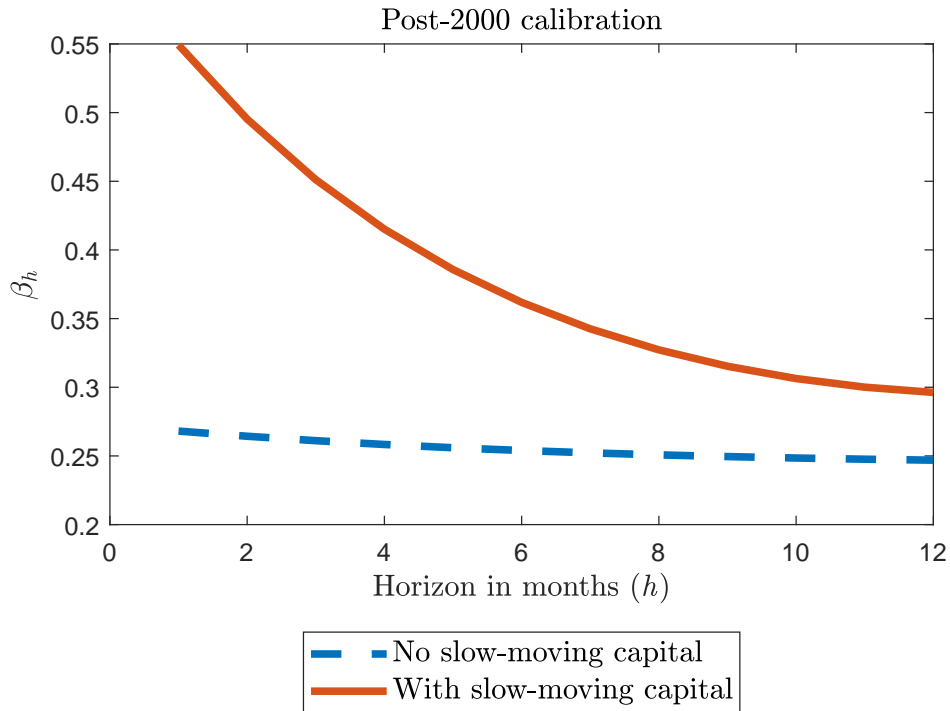
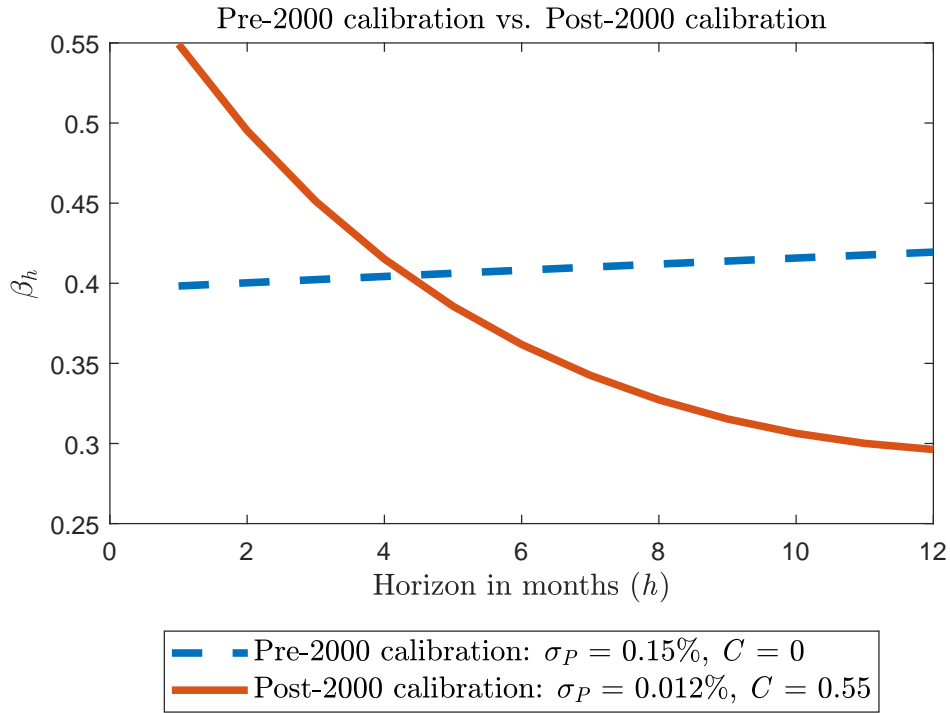


FIGURE V
Model-implied coefficients β_h versus horizon (h) in months

The first figure shows the model-implied β_h coefficients from equation (21) for the pre-2000 and post-2000 calibrations discussed in the text. The second figure isolates the role of slow-moving capital in the post-2000 calibration, alternately setting $q = 100\%$ (“No slow-moving capital”) and $q = 30\%$ (“With slow-moving capital”).