Imperfect Inflation Expectations and Treasury Bond Returns*

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Abstract

This study examines how imperfect inflation expectations affect excess returns on Treasury bonds. Survey forecasts and market responses to inflation announcements indicate that investors underreact to inflation news. Consistent with this slow updating of beliefs, bond excess returns are predictable: excess returns are significantly lower following upward revisions in inflation forecasts. I incorporate sticky inflation beliefs into a term structure model that explains both this predictability and the rejection of the expectations hypothesis without time-varying bond risk premia. Shocks to expected inflation also explain a smaller share of nominal yield movements than standard macro-finance models predict.

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1 Introduction

Inflation expectations are central to asset pricing and monetary economics. Full-information rational expectations (FIRE), the workhorse assumption in macro-finance term structure models, implies that inflation news is incorporated into bond yields immediately and fully. Survey evidence, however, tells a different story. Households, firms, and professional forecasters systematically underreact to inflation news (Coibion and Gorodnichenko, 2012, 2015), and the post-pandemic inflation surge has reinforced this evidence (Koch and Noureldin, 2024).[1] If investors who trade Treasury bonds are subject to the same friction, what are the implications for the market most directly exposed to inflation risk? This paper addresses this question. I show that underreaction to inflation news leads bond yields to adjust gradually, thereby generating predictable bond excess returns.

This study begins by demonstrating that real-time inflation expectations exhibit systematic underreaction to recent inflation news. Using forecasts from the Survey of Professional Forecasters (SPF) spanning 1968 to 2024, this study confirms the stickiness of inflation forecasts: when forecasters revise their inflation expectations upward (downward), subsequent realizations systematically exceed (fall below) their forecasts, indicating that their revisions do not fully incorporate new information. Because survey respondents may not represent the marginal investor, I complement the surveys with an event study of Treasury yields around CPI announcements. If yields impounded inflation news on impact, the response of yield spreads would be complete on the announcement day; instead, yield spreads continue to drift in the direction of the surprise for roughly three months. Underreaction, in other words, is visible both in what forecasters say and in what market prices do.

If inflation news is slowly incorporated into bond yields, this pattern of underreaction translates into predictable patterns in realized bond excess returns. When positive inflation news arrives and investors do not fully adjust their expectations, bond yields fail to immediately reflect the full implications of this news. As investors' inflation beliefs gradually adjust, bond yields rise to incorporate the information, generating predictably low returns. Following the literature (Duffee, 2018; Pang, 2026), I use revisions in SPF forecasts of the current-quarter inflation as a proxy for inflation news and find that bond excess returns remain low for several quarters following high inflation news. An econometrician assuming

[1] Sticky inflation expectations may arise from inattention (Mankiw and Reis, 2002) or noisy information (Sims, 2003). As these mechanisms generate similar belief dynamics in this setting, I abstract from their underlying source and focus on the asset-pricing implications of stickiness.

FIRE would observe a negative relation between inflation news and future bond excess returns. The magnitude is substantial: a one-standard-deviation increase in inflation news lowers the two-year bond's subsequent excess return by 23 basis points per annum (t-statistic = 3.29). Consistent results are obtained for bonds across alternative maturities, with an R^2 of 7%–8%, which is economically sizable. Furthermore, the predictive power of inflation news survives controls for well-known bond return predictors, extends to a household-survey measure of inflation news, and holds out of sample.

The negative slope coefficient on inflation news in bond return regressions is difficult to reconcile with risk-based explanations. In particular, the correlation between inflation news and inflation disagreement is positive (0.23) in this study's sample, suggesting that investors may face greater inflation uncertainty when positive inflation news arrives. Within a risk-based framework, such uncertainty commands a positive nominal term premium (Bansal and Shaliastovich, 2013). This prediction contradicts the negative relation between inflation news and future bond excess returns observed in the data.

Instead, this predictability in bond excess returns likely stems from systematic errors in inflation expectations rather than time-varying risk premia. To support this interpretation, I control for various well-known bond factors in the predictive regression and find that the predictive power of inflation news remains robust. Additionally, I construct a measure of the wedge between the Fed and SPF inflation forecasts that strongly predicts future SPF inflation forecast errors. This measure predicts future bond excess returns with an R^2 of 10%–11%, with predictive power likely stemming from Fed staff's efficient use of public information (Bauer and Swanson, 2023). Furthermore, the predictive power of inflation news for future bond excess returns is time-varying, strengthening when survey beliefs about inflation are stickier.

Simultaneously, sticky inflation beliefs help explain the variance decomposition of nominal yield shocks. Duffee (2018) revisits the importance of inflation news in Treasury bond yield innovations and suggests that existing models overweight inflation news relative to reality. Specifically, the variance of news about expected inflation accounts for only 10%–20% of the variance of yield shocks at a quarterly frequency. This is much lower than standard macro-finance term structure models imply (Piazzesi and Schneider, 2007; Bansal and Shaliastovich, 2013; Kung, 2015), which typically generate variance ratios of approximately 100%. At first glance, this finding appears difficult to reconcile with the bond return predictability I document: if inflation news strongly predicts future bond excess returns, how can it account for so little of yield variance? This study shows that sticky inflation expectations

complement the existing explanations of this puzzle. When investors underreact to inflation news, the contemporaneous response of yields to inflation shocks is attenuated, reducing the contribution of inflation news to overall nominal yield variance.

To provide a theoretical foundation for this study's empirical findings, I develop a consumption-based asset pricing model following Bansal and Shaliastovich (2013) and allow for biased beliefs about inflation dynamics. The representative agent in this model forms expectations about inflation that are "sticky" according to Mankiw and Reis (2002) and Coibion and Gorodnichenko (2015)—expectations adjust gradually to new information rather than fully incorporating available data. In the model, when positive inflation news arrives, the representative investor updates expectations by less than the FIRE benchmark would prescribe. Therefore, current yields do not fully incorporate the inflation news. As investors realize their bias and revise their inflation expectations upward, yields rise and realized excess returns on bonds are low. The converse holds for negative inflation news. This mechanism implies a negative relation between inflation news and future bond excess returns, and a positive relation between yields and future bond excess returns from the perspective of an econometrician, explaining the bond return predictability documented in Cochrane and Piazzesi (2005) and the failure of the expectations hypothesis for bond yields. Overall, the model generates several key predictions that match the empirical facts: (i) inflation forecast revisions positively predict realized forecast errors; (ii) inflation news negatively predicts future bond excess returns; (iii) yields positively predict future bond excess returns; and (iv) the inflation variance ratio is attenuated relative to that of standard macro-finance term structure models.

This study contributes to the growing body of research in finance and macroeconomics that emphasizes the role of expectations formation by relaxing the FIRE assumption (Mankiw and Reis, 2002; Bordalo et al., 2020). Using survey data, Coibion and Gorodnichenko (2015) provide empirical support for such belief frictions and document that professional forecasters underreact to macroeconomic news in a manner consistent with models of sticky beliefs or noisy signals. Several studies analyze the implications of such frictions for expectations of monetary policy (Bauer, Pflueger, and Sunderam, 2024; Schrimpf and Sihvonen, 2025), interest rates (Cieslak, 2018; Wang, 2021), and recessions (Gómez-Cram, 2022; Su and Yu, 2025). Closely related to my work, Cieslak (2018) finds that biased short-rate expectations generate Treasury bond return predictability, and that these expectation errors are primarily driven by *real rate* rather than inflation shocks. I show that underreaction in inflation expectations also generates bond return predictability

and accounts for key puzzles in the bond market, and that this predictability is distinct from the biased real rate expectations channel documented by Cieslak (2018). Pang (2026) finds that long-term inflation expectations are less sticky and more extrapolative than short-term expectations, explaining the excess sensitivity of yields to news at the long end. Delao and Myers (2024) find that long-term inflation survey forecasts explain *contemporaneous* bond yield movements better than realized inflation does. My work builds on these studies but differs in emphasizing the effects of sticky inflation expectations on bond excess return predictability and the attenuated response of yields to inflation news across maturities. Relatedly, Hong, Sraer, and Yu (2017) show that inflation disagreement, combined with short-sale constraints, generates overpricing of long-term bonds. Their channel operates through the dispersion of beliefs, whereas mine operates through the sluggish adjustment of the consensus.

This study offers important insights into the predictability of Treasury bond risk premia. The recent empirical literature documents several predictors of Treasury bond excess returns, including forward rates (Cochrane and Piazzesi, 2005), interest rate cycles (Cieslak and Povala, 2015), macro fundamentals (Ludvigson and Ng, 2009), output gap (Cooper and Priestley, 2009), inflation and real uncertainty (Bansal and Shaliastovich, 2013), and inflation disagreement (Hong, Sraer, and Yu, 2017). This study extends this line of research by showing that inflation news can also predict bond excess returns, with clear theoretical foundations. Unlike predictors estimated as linear combinations of forward rates, inflation news is observable in real time and requires no parameter estimation, avoiding generated-regressor and in-sample overfitting concerns associated with yield-based factors.

Finally, this study contributes to the spanning controversy in the macro-finance literature—the debate over whether macro-based return predictors are spanned by bond yields. Recent empirical evidence suggests that the yield curve does not summarize all predictive information about future Treasury bond returns (Ludvigson and Ng, 2009). As a possible solution, Duffee (2011) and Joslin, Priebsch, and Singleton (2014) propose models in which an unspanned factor moves risk premia and expected future short rates in offsetting directions, leaving the current term structure unaffected. Cieslak (2018) finds that investors' expectations of the real short rate systematically differ from ex-post realizations, leading to predictability in bond excess returns beyond that captured by yields. Bianchi, Büchner, and Tamoni (2021) and Huang and Shi (2023) identify macro factors that predict future bond returns by applying machine learning regularization techniques to a comprehensive set of real-time macroeconomic variables. This study suggests that, when investors underreact to

inflation news, the information contained in recent inflation news is not fully incorporated into current yields, leaving room for inflation variables to forecast future bond returns.

The remainder of this paper is organized as follows: Section 2 describes the data. Section 3 provides evidence on sticky inflation expectations using announcement and survey data. Section 4 presents the bond excess return predictability from inflation news. Section 5 develops a dynamic model with sticky inflation beliefs. Finally, Section 6 concludes.

2 Data and Variables

This section describes the data sources, the construction of Treasury bond excess returns, and the key predictors used in the empirical analysis.

2.1 Treasury bond returns

The empirical analysis uses daily Treasury yield data from Gürkaynak, Sack, and Wright (2007) to construct annual excess returns for zero-coupon Treasury bonds from 1968 to 2024 (yield data from 2025 are used to calculate annual returns).[2] The Gürkaynak, Sack, and Wright (2007) dataset provides continuously compounded zero-coupon yields estimated from a smooth yield curve fitted to off-the-run Treasury securities.[3] As the SPF forecasts are typically conducted in the second month of each quarter, I sample yields at the end of that month and construct holding period returns for zero-coupon bonds. This timing ensures that using the SPF forecasts as predictors does not introduce forward-looking bias.

Based on the yields, the realized one-year excess return on a zero-coupon bond with n periods to maturity is defined as:

$$rx^{\$}_{t\to t+1,n} = ny^{\$}_{t,n} - (n-1)y^{\$}_{t+1,n-1} - y^{\$}_{t,1}, \tag{1}$$

where $y^{\$}_{t,n}$ is the n-period nominal yield at time t. A dollar superscript is used to denote nominal quantities. Returns and yields are expressed in percent per annum. This study focuses on bond maturities of $n = 2, 3, 5, 10$ years for ease of exposition. Panels A and B of

[2]The data are available at `https://www.federalreserve.gov/data/nominal-yield-curve.htm`.

[3]In untabulated results, I verify robustness using the Center for Research in Security Prices database (Fama Risk-Free Rates and Fama-Bliss Discount Bond Yields), which contains prices for 1- through 5-year zero-coupon Treasury bonds.

Table 1 presents summary statistics for bond excess returns with a one-year holding period. Panel C reports contemporaneous correlations across bond maturities, confirming that bond excess returns are strongly correlated across maturities.

2.2 Inflation expectations

Following Pennacchi (1991), Duffee (2018), and Cascaldi-Garcia (2025), I infer inflation expectations from surveys of market practitioners, as consensus forecasts (cross-sectional means) from these surveys may proxy for the beliefs of sophisticated market participants. Specifically, I use forecasts of the GDP deflator from the SPF maintained by the Federal Reserve Bank of Philadelphia. The SPF forecast term structure spans from the nowcast (current quarter) to four quarters ahead. Following Bordalo et al. (2020), I construct consensus forecasts by averaging individual forecasts after excluding, for each horizon and quarter, observations that lie more than five interquartile ranges from the cross-sectional median. As in Duffee (2018), I use revisions in current-quarter inflation forecasts (nowcasts) to measure inflation news in the main analysis. Robustness tests consider forecast revisions at longer horizons and alternative measures of inflation expectations and news.

The top panel of Figure 1 plots the inflation rate and term structure of the SPF inflation forecasts. The visual impression from the graph suggests that survey forecasts move within a narrower band than realized inflation, appearing sluggish in responding to inflation news. For example, between 2020 and 2021, inflation forecasts fall substantially below realizations, showing underreaction to the sharp rise in inflation. The bottom panel of Figure 1 plots inflation news, measured by SPF forecast revisions for current-quarter inflation, alongside excess returns on the five-year bond. During this recent period, revisions to inflation forecasts are positive and large, while excess returns on long-term bonds are persistently negative, suggesting that investors do not fully anticipate the inflation increase.

In contrast, the Fed's inflation expectations are closer to FIRE than those of the SPF. Thus, I use the Federal Reserve Greenbook forecasts, prepared by the staff of the Board of Governors ahead of each Federal Open Market Committee (FOMC) meeting, as an arguably rational benchmark for inflation expectations. The Greenbook forecasts are published on the Federal Reserve Bank of Philadelphia website with a five-year lag and are available from 1965Q4 to 2020Q4. The Greenbook forecasts inflation (as measured by the GDP deflator) for five or six quarters ahead, although the forecast horizon varies over time and with the date of FOMC meetings. I retain the last Greenbook forecast released in the first two months

of a quarter to align the Greenbook and SPF forecasters' information sets.

2.3 Predictor variables and event study

In the predictive regressions, I include as control variables the standard bond factors and macro time series commonly used in the literature. Specifically, I consider the (1) Cochrane and Piazzesi (2005) yield factor CP, (2) Cieslak and Povala (2015) cycle factor CF, (3) Cieslak and Povala (2015) trend inflation τ^{CPI}, (4) Cooper and Priestley (2009) output gap factor GAP, (5) 3-month Treasury bill rate TBL, (6) inflation disagreement $Inflation_vol$ (Bansal and Shaliastovich, 2013; Hong, Sraer, and Yu, 2017), and (7) real GDP growth disagreement $Real_vol$.[4] For robustness, I also include the (8) modified Ludvigson and Ng (2011) factor LN, (9) hidden component H of the one-dimensional term premium factor constructed by Duffee (2011), and (10) Berardi et al. (2021) convergence gap CG. These three series are obtained from Huang and Shi (2023) and have shorter sample periods owing to data availability and are thus used in robustness tests.

Finally, I examine daily bond yield spread responses to surprises in core Consumer Price Index (CPI) announcements in an event study. For this analysis, survey forecasts of core CPI are obtained from Money Market Services (MMS) and are available from 1989 to 2025. Following Gürkaynak, Sack, and Swanson (2005), I measure market expectations using the median forecast compiled and published by MMS on the Friday before each release. Historical data on the month-over-month percentage change in core CPI for urban consumers (excluding food and energy) are published monthly by the Bureau of Labor Statistics. The surprise is defined as the difference between the realization and the forecast.

3 Evidence of Sticky Inflation Expectations

3.1 Do forecast revisions predict forecast errors?

Traditional macroeconomic and finance models often assume that agents form expectations according to FIRE. Under FIRE, agents use all available information to form the most accurate forecasts. A key implication of FIRE is that one cannot predict future

[4]The forecast disagreement measures for inflation and real GDP growth are defined as the differences between the 75th and 25th percentiles of the SPF panelists' one-quarter-ahead forecasts.

realized forecast errors based on any information available today. To assess this possibility, the top panel of Figure 1 plots the term structure of inflation forecasts over time, alongside realized current-quarter inflation. The forecast horizon ranges from the current quarter to four quarters ahead, and the vertical distance between each forecast point and the solid line indicates the magnitude of the corresponding forecast error. The visual impression from the graph suggests that realizations tend to exceed forecasts during inflation surges. In contrast, realizations tend to fall below forecasts following inflation drops. Additionally, survey forecasts of inflation move within a narrower range than realized inflation and are often inadequate, a pattern inconsistent with FIRE but consistent with sticky beliefs.

To formally investigate this issue, following Coibion and Gorodnichenko (2015), I test FIRE using the following regression specification:

$$\pi_{t+\frac{h}{4}} - \tilde{E}_t(\pi_{t+\frac{h}{4}}) = \alpha + \beta \left(\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}}) \right) + \epsilon_{t,t+\frac{h}{4}}, \tag{2}$$

where $\tilde{E}_t(\cdot)$ denotes the subjective expectation operator at time t, $\pi_{t+\frac{h}{4}} - \tilde{E}_t(\pi_{t+\frac{h}{4}})$ is the h-quarter-ahead inflation (GDP deflator) forecast error and $\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}})$ is the inflation forecast revision. This regression is derived from the immediate implication of FIRE that past forecast revisions, which are available when the forecasts are made, should not be correlated with future forecast errors, implying that β should be indistinguishable from 0. A positive coefficient β implies that expectations are sticky: when the average forecast revision is positive (negative), the consensus forecast is not sufficiently high (low).

Table 2 reports results for an econometric test of predictability. Columns (1)-(3) show that revisions in inflation expectations positively predict subsequent forecast errors. For example, at the three-quarter horizon, the estimated coefficient is 1.57, with a t-statistic of 3.54, close to the estimate in Bordalo et al. (2020) from a shorter sample. I next examine the predictability of inflation forecast revisions, which should be serially uncorrelated under FIRE. Columns (4)-(6), however, show that forecast revisions are positively autocorrelated: forecasters tend to revise their expectations in the same direction in consecutive quarters. This pattern is consistent with an insufficient initial response to inflation news. These tests should be interpreted as tests of the joint hypothesis of full information and rational expectations. A rejection of the FIRE restrictions may therefore reflect either departures from rational expectations or imperfect information about inflation. Overall, the regression results reinforce the main message of Figure 1: inflation expectations adjust only gradually to recent news. The findings thus reject FIRE and are consistent with existing evidence of

underreaction in the inflation expectations of both households and professional forecasters (Coibion and Gorodnichenko, 2012, 2015).

3.2 Evidence from inflation surprises

One concern is that survey forecasts may not perfectly reflect the beliefs of market participants. To mitigate this concern, I complement the survey evidence by examining the response of bond yields to inflation surprises. Following Cieslak and Povala (2015) and Gürkaynak, Sack, and Swanson (2005), I use core CPI, which excludes food and energy and is closely watched by bond investors and policymakers. To examine the effects of inflation surprises, I follow Jordà (2005) and Swanson (2021) and run a series of regressions at multiple horizons, indexed by h, of the following form:

$$\tilde{y}^{\$}_{\tau-1+h,n} = \alpha_h + \beta_h \tilde{y}^{\$}_{\tau-1,n} + \gamma_h \tilde{S}_\tau + \epsilon^h_\tau, \tag{3}$$

where τ denotes the date of a CPI release, $\tau - 1$ denotes the business day before the release, $\tilde{y}^{\$}_{\tau,n}$ denotes the yield spread between n-year and 1-year bonds at the close of day τ, $\tilde{S}_\tau$ denotes the core CPI surprise (realized minus MMS forecast) on release date τ, and h denotes the forecast horizon measured in business days, with each horizon associated with its own regression. The coefficient of interest is γ_h.

I estimate that the coefficients α_h and β_h are close to zero and one, respectively; therefore I impose those restrictions in the analysis that follows. Equation (3) then becomes a regression of the h-day change, $\tilde{y}^{\$}_{\tau-1+h,n} - \tilde{y}^{\$}_{\tau-1,n}$, on the core CPI surprise $\tilde{S}_\tau$ for each CPI release date τ. The estimated coefficients $\hat{\gamma}_h$ vary across forecast horizons h, allowing me to plot them as a function of the horizon h to examine whether they rise as h increases. For longer horizons h, a greater amount of non-inflation news affects yields; therefore the residuals ϵ^h_τ and standard errors of the coefficient estimates $\hat{\gamma}_h$ are larger.

Figure 2 plots the results of these regressions for the effects of the core CPI surprise on yield spreads. The solid blue line in each panel plots the point estimates of $\hat{\gamma}_h$ as a function of horizon h, and the dashed orange lines plot 90% confidence intervals around the point estimates. Yield spreads jump on impact at $h = 0$, but the initial response accounts for only part of the subsequent adjustment. A one-standard-deviation core CPI surprise raises the 2-year yield spread by about 0.5 bp on announcement day, with the cumulative effect rising to roughly 0.9 bp after 30 business days and 1.3 bp after 60 business days. Under FIRE, the

drift should be zero; instead, less than half of the eventual response occurs on impact. These findings suggest that inflation news is slowly impounded into yields, consistent with sticky inflation expectations. Thus, the collective evidence from survey forecasts and inflation announcements indicates stickiness in inflation expectations.

4 Inflation News and Bond Returns

4.1 In-sample evidence

Having established that inflation expectations exhibit systematic stickiness, I now turn to the central question: does this underreaction generate predictable patterns in Treasury bond excess returns? The intuition is straightforward. When positive inflation news arrives but investors fail to fully adjust their expectations, bond yields do not immediately reflect the full implications of the news. As beliefs gradually correct, yields rise to incorporate the information that was initially underweighted, and bondholders earn predictably low returns. The converse holds for negative inflation news. This mechanism implies a negative relation between inflation news and subsequent bond excess returns.

To test this prediction, I follow Duffee (2018) and use revisions in current-quarter inflation forecasts (nowcasts) from the SPF as a measure of inflation news. Specifically, I estimate the following predictive regression for each bond maturity n:

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1}, \tag{4}$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of the current-quarter inflation rate, and V_t is the vector of control variables as described in Section 2. The independent variables are standardized to have zero mean and unit variance to facilitate interpretation.

Table 3 reports the results of univariate predictive regressions with inflation news as the predictor. Column (1) of Panel A shows that revisions in expectations of the inflation rate negatively forecast subsequent bond excess returns. For the 2-year bond, the estimated coefficient is -0.23, with a t-statistic of -3.29.[5] This implies that a one-standard-deviation

[5]The Stambaugh (1999) small-sample bias increases with predictor persistence. Inflation news has an autocorrelation of 0.26, well below that of yield-based predictors, making this bias less concerning.

increase in inflation news is associated with a 23 bps lower excess return on a 2-year bond, which is economically sizable. This result is consistent with the explanation that investors underreact to inflation news and slowly update their beliefs about inflation, producing predictable bond returns. The remaining columns of Panel A show that the predictive power persists for bonds with other maturities. Moreover, as shown in the bottom row of Panel A, inflation news has R^2 values of 7%–8%.

The negative relation is difficult to reconcile with risk-based explanations. In particular, the correlation between inflation news and SPF inflation disagreement is positive (0.23) in my sample, suggesting that investors may face greater inflation uncertainty when positive inflation news arrives. Yet under a risk-based framework, such uncertainty would command a positive term premium, as argued by Bansal and Shaliastovich (2013). This prediction stands in contrast to the pattern observed in the data.

However, one may be concerned that the predictive power of inflation news is confounded by other factors. To mitigate this possibility, I include various control variables in Panel B, including the Cochrane and Piazzesi (2005) yield factor, Cieslak and Povala (2015) cycle factor, and uncertainty about future real growth and inflation (Bansal and Shaliastovich, 2013). After directly controlling for these factors, the predictive ability of inflation news is not materially affected. As shown in column (1) of Panel B, a one-standard-deviation increase in inflation news is associated with a 22 bps (t-statistic = 3.87) lower excess return on a 2-year bond, which is slightly lower than that from the univariate regression. Furthermore, I consider additional bond factors and report the results in Table A3 in the Appendix for robustness. These factors, such as the modified Ludvigson and Ng (2011) factor LN_t, hidden component H_t of the Duffee (2011) one-dimensional term premium factor, and Berardi et al. (2021) convergence gap CG_t, have shorter sample periods because of data availability. Table A3 shows that the coefficients on inflation news remain statistically significant at the 1% level. Thus, inflation news exhibits incremental predictive power for future bond excess returns beyond various bond factors, including bond yields and inflation uncertainty.

Another concern is that annual bond excess returns are sampled quarterly and thus overlap, potentially causing Newey-West t-statistics to overstate significance in long-horizon predictive regressions (Wei and Wright, 2013). Table A1 therefore repeats the tests in Table 3 using the reverse-regression approach of Wei and Wright (2013), which corrects for overlapping returns. Inflation news remains a significant predictor across maturities under this more conservative inference.

I also provide direct evidence that inflation forecast errors are indeed driving the predictive power of inflation news. Specifically, Table A2 shows that after controlling for SPF inflation forecast errors, the predictive power of inflation news for future bond excess returns is largely subsumed. For instance, a one-standard-deviation increase in inflation news is associated with only an 8 bps decrease (t-statistic = 1.49) in future two-year bond excess returns after controlling for inflation forecast errors. This finding suggests that the predictive power of inflation news likely stems from systematic bias in inflation expectation errors, leading to predictable returns as investors subsequently update their beliefs.

Inflation news predicts realized bond excess returns with declining strength across horizons as investors update their beliefs. To illustrate this, Figure 3 plots the coefficients from regressing annual bond excess returns on inflation news at horizons ranging from 0 to 8 quarters. As can be seen from the figure, bond excess returns react negatively to inflation news during the first four quarters after such news arrives; however, the effect gradually diminishes through 8 quarters, consistent with the observation that investors and bond prices gradually incorporate past inflation news.

Thus far, the analysis uses inflation news for the current quarter as the predictor. I present robustness tests to show that the results hold for inflation news at longer horizons. To explore this, Figure A1 in the Appendix plots the relation between inflation news across alternative horizons and annual bond excess returns. As shown in Figure A1, the predictive power of the inflation forecast revisions for 1- to 3-quarter horizons is statistically significant and lines up closely with the results based on forecast revisions for the current quarter in Table 3.[6] This aligns with the observation that inflation news is highly correlated across horizons.

4.2 Evidence from alternative expectation measures

The evidence so far is based on professional forecasters' inflation expectations. If bond return predictability reflects the slow incorporation of publicly available inflation news, similar patterns should emerge in the inflation news reaching households, which hold a non-negligible share of Treasury securities.[7] To examine this possibility, I construct a survey-

[6]Forecast revisions can be computed only up to 3 quarters ahead, because the 3-quarter-ahead revision uses the 4-quarter-ahead forecast issued in the previous quarter, and 4 quarters is the longest horizon in the SPF's term structure of quarterly forecasts.

[7]According to the Federal Reserve's Financial Accounts, households and nonprofit organizations held approximately 10% of outstanding Treasury securities at the end of 2025.

based measure of household inflation news from the Michigan Survey of Consumers. The survey asks respondents whether they have heard of any favorable or unfavorable changes in business conditions, and responses mentioning prices are coded as unfavorable (higher prices) or favorable (lower prices). I define $Household_news_t$ as the net percentage of respondents hearing unfavorable price news, that is, the percentage reporting unfavorable news about higher prices minus the percentage reporting favorable news about lower prices. Figure A2 in the Appendix plots the series, with construction details provided in the caption. Net unfavorable price news rises sharply during the 1979-1982 inflation episode and the 2021-2022 inflation surge. Because it captures household inflation news rather than professional forecasts, this measure provides an independent test of the predictive power of inflation news.

Table 4 shows that household inflation news negatively predicts bond excess returns, mirroring the SPF-based evidence in Table 3. In Panel A, a one-standard-deviation increase in net unfavorable inflation news is associated with a 26 bps lower excess return on the 2-year bond (t-statistic = -2.66). The estimated coefficients are stable across maturities, ranging from -0.26 to -0.31, with predictive R^2s of 8%–9%, comparable to those obtained using SPF inflation news. Panel B shows that the predictive power is essentially unaffected by including the bond and macro controls: the coefficient for the 2-year bond is -0.25 (t-statistic = -2.54), and the loadings remain statistically significant at all maturities, reaching -0.29 (t-statistic = -3.14) for the 10-year bond. Thus, household inflation news predicts bond returns with the same sign and a similar magnitude as professional forecasts.

The evidence of underreaction raises a key question: is it possible to use real-time information more efficiently than survey forecasters do when forming inflation forecasts? Addressing this requires a proxy for objective inflation expectations in real time. The literature (e.g., Romer and Romer, 2000) documents that the Greenbook forecasts, prepared by Fed staff, predict inflation more accurately than private-sector forecasts, including those from the SPF. This superior accuracy could reflect either more efficient use of information by Fed staff or their access to additional information. However, Bauer and Swanson (2023) find that Fed forecasts are simply responding to public news, supporting the former explanation. The difference between Greenbook and SPF inflation forecasts can therefore be used to capture bias in SPF inflation expectations. To measure this difference, I define the "wedge" variable as follows:

$$\Delta_t(\pi_{t+1}) \equiv E_t[\pi_{t+1}] - \tilde{E}_t[\pi_{t+1}]. \tag{5}$$

Here, $E_t[\pi_{t+1}]$ and $\tilde{E}_t[\pi_{t+1}]$ denote the Greenbook and SPF one-year-ahead inflation forecasts, respectively.

The top panel of Figure 4 depicts the time variation in the Greenbook and SPF forecasts and realized inflation. The graph suggests that the SPF expectations systematically lag behind those of the Greenbook. For example, as the bottom panel shows, the wedge in inflation forecasts is high during the stagflation of the 1970s. Table A4 confirms that the Greenbook forecasts outperform SPF forecasts in predicting future inflation. When both forecasts are included in a horse-race regression, the Greenbook absorbs the predictive power of the SPF forecast. Moreover, the wedge between the Greenbook and SPF forecasts significantly and positively predicts SPF forecast errors (t-statistic = 3.21), confirming the superior performance of the Greenbook forecasts.

Panel A of Table 5 shows that the wedge between the Greenbook and SPF inflation forecasts negatively predicts future bond excess returns. Moreover, the economic magnitude is large: a one-standard-deviation increase in the wedge variable decreases the 2-year bond excess return by 28 bps (t-statistic = 2.42). Similar effects are also observed for bonds with alternative maturities. The R^2s in these regressions are 10%–11%, which are similar to those obtained using inflation news as a predictor. I add various control variables in Panel B, finding that the wedge variable still negatively predicts bond excess returns, albeit with generally smaller statistical significance. For the 2-year bond, a one-standard-deviation increase in the wedge variable is now associated with a 24 bps (t-statistic = 2.20) decrease in the 2-year bond excess return. In sum, the wedge between Fed staff and professional forecasters' inflation forecasts contains information about future bond excess returns, likely due to capturing inflation forecast biases of professional forecasters and market participants alike.

4.3 Biased short-rate expectations?

A critical concern is that investors may also have biased expectations of the short-term interest rate, which could confound the observed return predictability. Cieslak (2018) documents that investors systematically overestimate the future short rate, particularly during economic downturns, and that forecast errors are primarily driven by real rate shocks rather than inflation. If such biased short-rate expectations also generate bond return predictability, the predictive power I attribute to inflation news could instead reflect biased updating of real rate expectations.

To address this concern, I follow Cieslak (2018) and construct a short-rate expectations wedge, defined as the difference between an econometric forecast of the 3-month Treasury bill

rate and the corresponding SPF forecast. The econometric forecast is estimated over a 5-year rolling window using lagged bill rates and employment growth, based only on information available through time t to ensure a fair comparison with the time-t SPF forecast. Including this wedge in the regression specified in Equation (4) controls for the component of bond return predictability that arises from biased short-rate expectations. As shown in Table 6, the coefficient on inflation news remains statistically significant, though its significance weakens somewhat. This result suggests that the predictive content of inflation news is not subsumed by errors in short-rate expectations.[8]

I further investigate whether real rate news directly predicts bond excess returns. Specifically, I test whether revisions in real rate forecasts, constructed analogously to the inflation news measure, have predictive power for future bond returns (Table 7). Panel A shows that real rate news has virtually no predictive power. Take the 2-year bond as an example: a one-standard-deviation increase in real rate news is associated with an increase of only 4 basis points in its subsequent excess return (t-statistic = 0.53). The R^2 values are also near zero. After controlling for well-known bond return predictors, real rate news remains economically small and at best marginally significant as a predictor of bond excess returns (Panel B), suggesting that it is unlikely to be the primary source of the bond return predictability documented in this paper.

Overall, the combined evidence suggests that inflation news negatively predicts future bond excess returns, consistent with the notion that investors are slow to update their inflation beliefs. Additionally, the predictive power of inflation news is not subsumed by the many well-known bond return predictors considered. To further validate this finding, I turn to the out-of-sample predictive power of inflation news in the analysis that follows.

4.4 Out-of-sample accuracy

Recent studies suggest that the predictive power of bond return predictors may not hold out of sample (Thornton and Valente, 2012). Therefore, I evaluate the out-of-sample forecasting performance of inflation news for bond excess returns to assess whether the in-sample predictability documented above remains robust out of sample. Following Cieslak and Povala (2015), I form recursive one-year-ahead forecasts where the parameters of the

[8] An alternative possibility is that inflation news signals future short-rate changes. To mitigate this concern, in untabulated results I control for the perceived monetary policy response to the inflation gap of Bauer, Pflueger, and Sunderam (2024) and find that this measure does not significantly predict future bond excess returns, while the inflation news measure does.

predictive model are estimated using only information available at the time of the forecast, specifically employing inflation news observed through quarter $t-4$ and returns realized through quarter t to predict excess returns at $t+4$.

The out-of-sample evaluation employs four complementary metrics. The first is the ENC-NEW encompassing test of Clark and McCracken (2001), which examines whether the restricted model based on historical average returns encompasses all relevant predictive information, or whether inflation news contributes additional forecasting power beyond what is already captured by the benchmark. A rejection of the null hypothesis indicates that inflation news provides incremental predictive content beyond the historical average forecast. The ENC-NEW test statistic for maturity n is given by

$$\text{ENC-NEW}^{(n)} = P\frac{\sum_{t=1}^{T}\left(u_{t+4}^{2,(n)} - u_{t+4}^{(n)}\varepsilon_{t+4}^{(n)}\right)}{\sum_{t=1}^{T}\varepsilon_{t+4}^{2,(n)}}, \tag{6}$$

where P is the number of 1-year-ahead predictions and $\varepsilon_t^{(n)}$ and $u_t^{(n)}$ denote the prediction errors from the unrestricted and restricted models, respectively.

The second statistic is the out-of-sample R^2 proposed by Campbell and Thompson (2008), which compares the mean-squared prediction errors from the inflation news model against those from a naive forecast based on the historical average excess return:

$$R_{oos}^{2,(n)} = 1 - \frac{\sum_{t=1}^{T-4}\left(rx_{t\to t+4,n}^{\$} - \widehat{rx}_{\pi,t\to t+4,n}^{\$}\right)^2}{\sum_{t=1}^{T-4}\left(rx_{t\to t+4,n}^{\$} - \overline{rx}_{t\to t+4,n}^{\$}\right)^2}, \tag{7}$$

where the time index t and sample size T are expressed in quarters. $\widehat{rx}_{\pi,t\to t+4,n}^{\$}$ is the forecast of the annual excess return based on time-t inflation news, where the parameters of the predictive model are estimated using inflation news up to time $t-4$ and returns realized up to time t. $\overline{rx}_{t\to t+4,n}^{\$}$ is the return forecast using the historical average excess return estimated through time t. A positive R_{oos}^2 indicates that the predictive model generates smaller forecast errors than the benchmark, with larger values reflecting greater forecasting gains. The third statistic is the MSE-F test of McCracken (2007), which provides a formal test of equal predictive accuracy between the competing models. Finally, I compute p-values following Clark and West (2007), which account for the nested model structure when comparing a predictor-based forecast to a historical average benchmark.

I start the out-of-sample exercise at two dates, 1978Q1 and 1985Q1, following Cieslak and

Povala (2015). Panel A of Table 8 demonstrates that inflation news generates economically meaningful and statistically significant out-of-sample predictability across bond maturities over the evaluation period of 1978Q1 to 2024Q4. The positive R^2_{oos} values indicate that conditioning on inflation news consistently reduces forecast errors relative to the historical average benchmark, confirming that the predictive relationship identified in sample extends to out-of-sample settings. The ENC-NEW statistics reject the null that the historical average model encompasses the inflation news predictor, while the MSE-F tests corroborate the finding of superior predictive accuracy. The associated p-values confirm that these improvements are statistically significant at conventional levels. As several studies (Joslin, Priebsch, and Singleton, 2014; Cieslak and Povala, 2015) focus on post-1984 samples, I repeat the analysis for the post-1984 sample and present the results in Panel B. The out-of-sample predictive power persists in this more recent sample. Together, these findings suggest that the predictive power of inflation news for bond excess returns persists out of sample.

4.5 Conditional bond return predictability

The proposed sticky inflation belief mechanism posits that investors should underreact more to inflation news when beliefs are stickier. To test this prediction, I exploit time variation in inflation belief stickiness. Belief stickiness may change over time when the underlying information structures change. For instance, consensus beliefs may become stickier when forecasters are less attentive or information is noisier (Coibion and Gorodnichenko, 2015; Bordalo et al., 2020).

Accordingly, I estimate time-varying belief stickiness ($\hat{\lambda}_t$) via rolling regressions on the SPF inflation forecasts. This measure is derived from the slope coefficient of a 10-year rolling regression of 3-quarter-ahead inflation forecast errors on forecast revisions, following the main specifications in Bordalo et al. (2020) and Coibion and Gorodnichenko (2015).[9] Subsequently, I run the regression in Equation (4) by further including inflation belief stickiness and its interaction with inflation news.

Table 9 presents the main findings. The sticky belief mechanism implies that a smaller fraction of inflation news is incorporated into current bond yields when belief stickiness is high, which in turn induces stronger predictability of future bond excess returns as beliefs

[9]In the context of sticky-information models, inflation belief stickiness is computed as $\hat{\lambda} = \hat{\beta}/(1+\hat{\beta})$, where $\hat{\beta}$ is the slope coefficient from the regression of forecast errors on forecast revisions.

converge to fundamentals.[10] Focusing on the 2-year bond in column (1), the coefficient estimate on the interaction term is negative (-0.14) and statistically significant, with a t-statistic of -3.79. In terms of economic magnitude, when belief stickiness is at its mean level, a one-standard-deviation increase in inflation news predicts a 32 bps decrease in future bond excess returns. However, when belief stickiness is one standard deviation above its mean, a one-standard-deviation increase in inflation news predicts a 46 bps decrease in future bond excess returns (0.32 + 0.14). Moreover, including the interaction term slightly increases the R^2 relative to that in Table 3. Similar results are obtained using bonds with alternative maturities. Thus, time variation in inflation belief stickiness helps explain the conditional bond return predictability I uncover.

5 Theoretical Framework

This section examines whether a single friction—sticky inflation expectations—can jointly account for the empirical findings documented above. Building on Bansal and Shaliastovich (2013) and Gómez-Cram and Yaron (2021), I develop a representative-agent endowment economy in which the agent forms sticky beliefs about inflation dynamics. The model admits analytical solutions and illustrates how inflation belief errors affect Treasury bond prices and expected excess returns, thereby providing a unified explanation for the evidence in Section 4. Detailed derivations of all results in this section are provided in Appendix B.

5.1 Preferences

The representative agent has preferences given by the recursive utility function of Epstein and Zin (1989):

$$U_t = \left[(1-\delta)C_t^{\frac{1-\gamma}{\theta}} + \delta\left[\tilde{E}_t\left(U_{t+1}^{1-\gamma}\right)\right]^{\frac{1}{\theta}}\right]^{\frac{\theta}{1-\gamma}}. \tag{8}$$

In this parameterization, $\tilde{E}_t(\cdot)$ denotes the conditional expectations under the agent's subjective beliefs; C_t is aggregate consumption; δ is the time discount factor; γ is the coefficient of relative risk aversion; and $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$, where ψ is the elasticity of intertemporal substitution (EIS). The parameters satisfy $\gamma, \psi > 1$, which is the standard preference parameter configuration for asset pricing models with Epstein-Zin preferences (Bansal and

[10] As shown in Equation (28) in Section 5, when belief stickiness is high, the link between inflation forecast revisions and forecast errors strengthens, and thus so does the link with future bond excess returns.

Yaron, 2004). As shown in Epstein and Zin (1989), the first-order condition for the representative agent's problem implies that the gross return $R_{i,t+1}$ on any tradable asset i satisfies the Euler equation:

$$1 = \tilde{E}_t[M_{t+1}R_{i,t+1}], \tag{9}$$

with a stochastic discount factor (SDF) M_{t+1} given by

$$m_{t+1} \equiv \ln M_{t+1} = \theta \ln \delta - \theta \frac{1}{\psi} \Delta c_{t+1} + (\theta - 1) r_{c,t+1}, \tag{10}$$

where $\Delta c_{t+1} = \ln(C_{t+1}/C_t)$ is the log growth rate of aggregate consumption and $r_{c,t+1}$ is the log return on the wealth portfolio, that is, the aggregate consumption claim that pays out aggregate consumption as dividends.

Therefore, the real return satisfies the following asset pricing equation:

$$\tilde{E}_t \left[\exp(m_{t+1} + r_{t+1})\right] = 1, \tag{11}$$

which holds for any continuously compounded return $r_{t+1} = \ln(R_{t+1})$, including that on the wealth portfolio. This allows one to express the SDF in Equation (10) in terms of the fundamental state variables and shocks in the economy, and to solve for equilibrium real asset prices. Similarly, the nominal discount factor that can be used to price nominal payoffs through the nominal version of the Euler equation in Equation (11) equals the real discount factor minus the inflation rate:

$$m^{\$}_{t+1} = m_{t+1} - \pi_{t+1}. \tag{12}$$

5.2 Economy dynamics and subjective beliefs

To solve for the equilibrium asset prices, I specify an exogenous inflation process π_{t+1} following Bansal and Shaliastovich (2013) and Piazzesi and Schneider (2007). I denote by $x_{c,t}$ and $x_{\pi,t}$ the expected consumption growth and expected inflation under objective expectations, respectively. The consumption and inflation dynamics are specified as follows:

$$\begin{aligned} \Delta c_{t+1} &= \mu_c + x_{c,t} + \sigma_c \eta_{c,t+1}, \\ \pi_{t+1} &= \mu_\pi + x_{\pi,t} + \sigma_\pi \eta_{\pi,t+1}, \end{aligned} \tag{13}$$

where $\eta_{c,t+1}$ and $\eta_{\pi,t+1}$ are the standard normal shocks and σ_c and σ_π are the volatilities of consumption and inflation, respectively. For parsimony and to isolate the effect of belief biases, I assume that these volatilities are constant. I model the dynamics of the expected growth and expected inflation factors as

$$\begin{aligned} x_{c,t+1} &= \rho_c x_{c,t} + \rho_{c\pi}\tilde{x}_{\pi,t} + \sigma_{xc}\eta_{xc,t+1}, \\ x_{\pi,t+1} &= \rho_\pi x_{\pi,t} + \sigma_{x\pi}\eta_{x\pi,t+1}, \end{aligned} \tag{14}$$

where $\eta_{xc,t+1}$ and $\eta_{x\pi,t+1}$ are the standard normal shocks and σ_{xc} and $\sigma_{x\pi}$ reflect the fluctuations in expected growth and expected inflation, respectively. The parameters ρ_c and ρ_π capture the persistence of expected real growth and expected inflation, respectively, and $\rho_{c\pi}$ reflects the non-neutral effect of expected inflation on future expected growth. As documented in the literature (Bansal and Shaliastovich, 2013; Piazzesi and Schneider, 2007), expected inflation negatively forecasts future real growth; accordingly, the inflation non-neutrality parameter $\rho_{c\pi}$ is negative.

The term $\tilde{x}_{\pi,t}$ in Equation (14) captures the representative agent's subjective inflation belief, which influences household consumption decisions through inflation non-neutrality. Motivated by the evidence on expected inflation from survey data in Section 3, I model the representative agent as holding sticky inflation beliefs. Consistent with models of inattention (Mankiw and Reis, 2002) or information rigidities (Sims, 2003), these beliefs follow

$$\begin{aligned} \tilde{E}_t[x_{\pi,t+1}] &= (1-\lambda)E_t[x_{\pi,t+1}] + \lambda\tilde{E}_{t-1}[x_{\pi,t+1}] \\ &= \rho_\pi \tilde{x}_{\pi,t}, \end{aligned} \tag{15}$$

where $\tilde{x}_{\pi,t} \equiv (1-\lambda)\sum_{j=0}^{\infty}(\lambda\rho_\pi)^j x_{\pi,t-j}$. The parameter $\lambda \in [0,1)$ governs the degree of belief stickiness: when $\lambda = 0$, inflation beliefs are rational and the model reduces to the constant-volatility setting of Bansal and Shaliastovich (2013); as λ increases toward one, beliefs become increasingly sluggish in responding to new information.

Under the representative agent's beliefs, the shock to expected inflation is

$$\begin{aligned} \sigma_{x\pi}\tilde{\eta}_{x\pi,t+1} &\equiv x_{\pi,t+1} - \tilde{E}_t[x_{\pi,t+1}] \\ &= \rho_\pi(x_{\pi,t} - \tilde{x}_{\pi,t}) + \sigma_{x\pi}\eta_{x\pi,t+1}, \end{aligned} \tag{16}$$

where $\rho_\pi(x_{\pi,t} - \tilde{x}_{\pi,t})$ represents the representative agent's inflation forecast bias. The sticky beliefs create a wedge between the agent's perceived inflation $\tilde{x}_{\pi,t}$ and rational

expectation $x_{\pi,t}$. As this bias is gradually realized by the representative agent, it leads to predictability of bond excess returns. Accordingly, the dynamics of expected inflation under the representative agent's beliefs can be rewritten as

$$\tilde{x}_{\pi,t+1} = \rho_\pi \tilde{x}_{\pi,t} + (1-\lambda)\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1}. \tag{17}$$

5.3 Bond pricing

In equilibrium, the price-consumption ratio is linear in the expected consumption and subjective inflation belief:

$$pc_t = A_0 + A_c x_{c,t} + A_\pi \tilde{x}_{\pi,t}, \tag{18}$$

where the equilibrium loadings and the constant are given by

$$\begin{aligned} A_c &= \frac{1-\frac{1}{\psi}}{1-\kappa_1\rho_c}, \qquad A_\pi = \frac{\kappa_1\rho_{c\pi}(1-\frac{1}{\psi})}{(1-\kappa_1\rho_\pi)(1-\kappa_1\rho_c)}, \\ A_0 &= \frac{\ln\delta + \kappa_0 + (1-\frac{1}{\psi})\mu_c + \frac{\theta}{2}(1-\frac{1}{\psi})^2\sigma_c^2 + \frac{\theta}{2}\kappa_1^2 A_c^2\sigma_{xc}^2 + \frac{\theta}{2}\kappa_1^2 A_\pi^2(1-\lambda)^2\sigma_{x\pi}^2}{1-\kappa_1}, \end{aligned} \tag{19}$$

where κ_0 and κ_1 are the linearization parameters from the Campbell-Shiller approximation. As in the standard long-run risks model of Bansal and Yaron (2004), when the intertemporal elasticity of substitution parameter ψ is greater than one, the intertemporal substitution effect dominates the wealth effect. In response to higher expected growth, agents demand more of the consumption claim, and consequently, the wealth-to-consumption ratio rises. Hence, the price-consumption loading on the expected consumption growth is positive, $A_c > 0$, and it is negative on the subjective expected inflation, $A_\pi < 0$, if inflation is bad for consumption ($\rho_{c\pi} < 0$).

Using the solution for the price-consumption ratio, I provide an analytical expression for the equilibrium SDF in terms of the fundamental state variables and shocks in the economy. Under the agent's subjective beliefs, the innovation in the SDF that determines the sources and compensations for risk in the economy is given by

$$m_{t+1} - \tilde{E}_t(m_{t+1}) = -\lambda_c\sigma_c\eta_{c,t+1} - \lambda_\pi\sigma_\pi\eta_{\pi,t+1} - \lambda_{xc}\sigma_{xc}\eta_{xc,t+1} - \lambda_{x\pi}\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1}. \tag{20}$$

The equilibrium market prices of short-run consumption and inflation risks and expected

consumption and expected inflation risks are equal to

$$\lambda_c = \gamma, \quad \lambda_\pi = 0, \quad \lambda_{xc} = (1-\theta)\kappa_1 A_c, \quad \lambda_{x\pi} = (1-\theta)(1-\lambda)\kappa_1 A_\pi. \tag{21}$$

The price of short-run consumption risk λ_c is equal to the risk-aversion coefficient γ, and as the immediate inflation innovations are independent of the real economy, their price of risk λ_π is zero. When agents have a preference for early resolution of uncertainty ($\gamma > 1/\psi$), the market price of expected consumption risk λ_{xc} is positive; the market price of expected inflation risk $\lambda_{x\pi}$ is negative if expected inflation is bad news for future consumption ($\rho_{c\pi} < 0$). Importantly, the market price of expected inflation risk is scaled by $(1-\lambda)$, reflecting that sticky beliefs attenuate the agent's perceived exposure to inflation news.

In the economy, the equilibrium solution to the nominal discount factor in Equation (12) satisfies

$$m^{\$}_{t+1} - \tilde{E}_t(m^{\$}_{t+1}) = -\lambda^{\$}_c \sigma_c \eta_{c,t+1} - \lambda^{\$}_\pi \sigma_\pi \eta_{\pi,t+1} - \lambda^{\$}_{xc} \sigma_{xc} \eta_{xc,t+1} - \lambda^{\$}_{x\pi} \sigma_{x\pi} \tilde{\eta}_{x\pi,t+1}, \tag{22}$$

where the nominal market price of the inflation risk $\lambda^{\$}_\pi$ is equal to one, and nominal market prices of expected growth and expected inflation risks are equal to their real counterparts.

Given the discount factor, the continuously compounded one-period real interest rate is affine in the state variables

$$r_{f,t} = -\tilde{E}_t(m_{t+1}) - \frac{1}{2}\tilde{Var}_t(m_{t+1}) = B_{0,1} + B_{c,1} x_{c,t} + B_{\pi,1} \tilde{x}_{\pi,t}. \tag{23}$$

This equation links variations in the real short rate to variations in fundamentals. Owing to the agent's intertemporal consumption-smoothing motive, the real short rate increases when one-period-ahead expected consumption growth rises (i.e., $B_{c,1} = \frac{1}{\psi} > 0$). In this case, the agent wishes to borrow to increase current consumption; however, as bonds are in zero net supply, aggregate borrowing cannot increase. In equilibrium, the real short rate must rise to induce the agent to borrow less.

Similarly, the one-period nominal interest rate is

$$r^{\$}_{f,t} = r_{f,t} + \tilde{E}_t(\pi_{t+1}) + \tilde{Cov}_t(m_{t+1}, \pi_{t+1}) - \frac{1}{2}\tilde{Var}_t(\pi_{t+1}). \tag{24}$$

The intuition underlying movements in the nominal short rate is straightforward. The first two terms on the right-hand side represent the standard Fisher equation. The third term

reflects that nominal rates are risky in real terms. If positive shocks to the inflation rate are associated with positive shocks to the SDF, the covariance term is positive and the agent requires additional compensation for holding the nominal short-term bond. Finally, the fourth term arises from Jensen's inequality.

Additionally, yields on zero-coupon bonds are affine functions of the state variables. Using the Euler equation, together with the real and nominal SDF, the n-period nominal yield is:

$$y^{\$}_{t,n} = \frac{1}{n}\left(B^{\$}_{0,n} + B^{\$}_{c,n}x_{c,t} + B^{\$}_{\pi,n}\tilde{x}_{\pi,t}\right), \tag{25}$$

where the loadings satisfy the recursions

$$\begin{aligned} B^{\$}_{c,n} &= \rho_c B^{\$}_{c,n-1} + \frac{1}{\psi}, \\ B^{\$}_{\pi,n} &= \rho_\pi B^{\$}_{\pi,n-1} + \rho_{c\pi} B^{\$}_{c,n-1} + 1, \end{aligned} \tag{26}$$

with initial conditions $B^{\$}_{c,0} = 0$ and $B^{\$}_{\pi,0} = 0$. Appendix B provides solutions for these loadings. In this model, the loading $B^{\$}_{c,n}$ is positive: nominal bond prices rise and yields fall following a negative shock to expected consumption growth. As higher inflation expectations decrease the valuations of nominal bonds, $B^{\$}_{\pi,n}$ is also positive.

5.4 Bond excess returns and the role of sticky beliefs

A key implication of sticky inflation beliefs is the distinction between expected returns under subjective and objective expectations. Under the agent's subjective beliefs, the one-period expected excess return on nominal bonds is determined by the negative covariation between innovations to excess log returns and innovations to the nominal SDF:

$$\begin{aligned} \tilde{E}_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}\tilde{Var}_t(rx^{\$}_{t\to t+1,n}) &= -\tilde{Cov}_t\left(m^{\$}_{t+1}, rx^{\$}_{t\to t+1,n}\right) \\ &= \underbrace{-\lambda^{\$}_{xc}B^{\$}_{c,n-1}\sigma^2_{xc}}_{\text{expected consumption risk}} \underbrace{-(1-\lambda)\lambda^{\$}_{x\pi}B^{\$}_{\pi,n-1}\sigma^2_{x\pi}}_{\text{expected inflation risk}}, \end{aligned} \tag{27}$$

where $\lambda^{\$}_{xc}$ and $\lambda^{\$}_{x\pi}$ denote the market prices of expected consumption and expected inflation risks, respectively. With a preference for early resolution of uncertainty, the equilibrium market price of expected consumption risk is positive ($\lambda_{xc} > 0$), that of expected inflation risk is negative ($\lambda_{x\pi} < 0$), and bond loadings on expected consumption and expected inflation

risks are both positive ($B^{\$}_{c,n-1} > 0, B^{\$}_{\pi,n-1} > 0$). This suggests that the nominal bond risk premium increases with inflation uncertainty but decreases with real uncertainty (Bansal and Shaliastovich, 2013).

Bond excess return predictability arises from biased inflation expectations because the expected excess return under the objective expectations (which is available to an econometrician) differs from that under the subjective expectations:

$$\begin{aligned} & E_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}Var_t(rx^{\$}_{t\to t+1,n}) \\ & = \tilde{E}_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}\tilde{Var}_t(rx^{\$}_{t\to t+1,n}) - B^{\$}_{\pi,n-1}\underbrace{\left[E_t(\tilde{x}_{\pi,t+1}) - \tilde{E}_t(\tilde{x}_{\pi,t+1})\right]}_{\text{inflation forecast bias}} \\ & = -\lambda^{\$}_{xc}B^{\$}_{c,n-1}\sigma^2_{xc} - (1-\lambda)\lambda^{\$}_{x\pi}B^{\$}_{\pi,n-1}\sigma^2_{x\pi} - (1-\lambda)\rho_\pi B^{\$}_{\pi,n-1}(x_{\pi,t} - \tilde{x}_{\pi,t}), \end{aligned} \tag{28}$$

where $E_t(\cdot)$ denotes conditional expectations under the objective beliefs. The additional term $(1-\lambda)\rho_\pi B^{\$}_{\pi,n-1}(x_{\pi,t} - \tilde{x}_{\pi,t})$ captures the predictability induced by belief biases. Following positive shocks to expected inflation, investors with sticky beliefs underestimate future inflation (i.e., $x_{\pi,t} > \tilde{x}_{\pi,t}$) and thus overvalue long-term bonds. As they gradually revise their inflation expectations upward, yields rise and bond excess returns become predictably low. As shown in Equation (28), recent inflation news positively predicts inflation forecast errors and, consequently, negatively predicts future bond excess returns, reconciling the findings in Tables 3 and 5. In contrast, without sticky inflation expectations (i.e., when $\lambda = 0$), $x_{\pi,t} = \tilde{x}_{\pi,t}$, and bond excess returns are unpredictable by past inflation news.

The expectations hypothesis posits that the risk premium on long-term bonds, or the expected excess return on long-term bonds over short-term bonds, should be constant over time. However, this has been consistently rejected by empirical evidence of time-varying risk premia (Cochrane and Piazzesi, 2005; Bansal and Shaliastovich, 2013). Equation (28) suggests that the model can account for both the predictability of bond returns using yield-based predictors, as documented by Cochrane and Piazzesi (2005), and the failure of the expectations hypothesis. Thus, the model generates apparent return predictability under the objective measure despite constant subjective risk compensation. In particular, when expected inflation is too high, yields also tend to be high in the model. As investors gradually recognize their forecast errors, they revise their inflation expectations downward, reducing bond yields and generating high excess returns as shown in Equation (1). Conversely, when expected inflation is too low, yields tend to be low, and subsequent upward revisions in inflation expectations lead to low excess returns. However, an econometrician assuming

FIRE would attribute the positive relation between yields and subsequent returns to a time-varying term premium, when in fact it reflects predictable errors in inflation expectations, a mechanism similar to that in Delao and Myers (2024).

5.5 Term premium

The nth period yield can be decomposed into expectations of average future short rates and average excess returns over the life of the bond:

$$y^{\$}_{t,n} = \frac{1}{n}\sum_{i=1}^{n} \tilde{E}_t y^{\$}_{t+i-1,1} + \frac{1}{n}\sum_{i=1}^{n} \tilde{E}_t rx^{\$}_{t+i-1\to t+i,n-i+1}. \tag{29}$$

The last term on the right is often described as the bond's nominal term premium, denoted as $tp^{\$}_{t,n}$. Given the assumption of constant conditional volatilities, the term premium is constant:

$$tp^{\$}_{t,n} = \frac{1}{n}B^{\$}_{0,n} - B^{\$}_{0,1}. \tag{30}$$

This result implies that, from investors' perspective, time variation in yields may reflect changing expectations of future short rates and inflation rather than time-varying risk premia. In the model, the apparent predictability of bond returns identified by an econometrician under rational expectations arises not from time-varying risk premia, but from systematic forecast errors induced by investors' sticky beliefs.

If I define the ex-ante real rate as the yield of a one-period nominal bond minus subjectively expected inflation, the nominal bond yield can be further decomposed as

$$y^{\$}_{t,n} = \frac{1}{n}\sum_{i=1}^{n} \tilde{E}_t r_{f,t+i-1} + \frac{1}{n}\sum_{i=1}^{n} \tilde{E}_t \pi_{t+i} + tp^{\$}_{t,n}. \tag{31}$$

Equation (31) shows that yield variations are driven by changes in expected future real rates and expected inflation, while the term premium contributes only to the level of yields.

5.6 Inflation variance ratios

This subsection analyzes what drives shocks to nominal bond yields. Following Duffee (2018), I define the unconditional inflation variance ratio of an nth period nominal bond as

the ratio of the variance of inflation news to the variance of yield shocks:

$$\text{Inflation Variance Ratio} \equiv \frac{Var(\epsilon^{(n)}_{\pi,t})}{Var(\epsilon^{(n)}_{y^\$,t})}, \tag{32}$$

where $\epsilon^{(n)}_{\pi,t}$ denotes the innovations to average expected inflation and $\epsilon^{(n)}_{y^\$,t}$ denotes the innovations to the n-period yield.

The model delivers a closed-form expression for the inflation variance ratio. Under sticky beliefs, innovations to average expected inflation are given by

$$\epsilon^{(n)}_{\pi,t} = \frac{1}{n}\frac{1-\rho_\pi^n}{1-\rho_\pi}(1-\lambda)(x_{\pi,t} - \rho_\pi \tilde{x}_{\pi,t-1}), \tag{33}$$

and innovations to the n-period yield are

$$\epsilon^{(n)}_{y^\$,t} = \frac{1}{n}B^\$_{c,n}\sigma_{xc}\eta_{xc,t} + \frac{1}{n}(1-\lambda)B^\$_{\pi,n}(x_{\pi,t} - \rho_\pi \tilde{x}_{\pi,t-1}). \tag{34}$$

As the term premium is constant under subjective expectations, shocks to the term premium are zero: $\epsilon^{(n)}_{tp^\$,t} = 0$.

Therefore, the model-implied inflation variance ratio is

$$\text{Inflation Variance Ratio} = \frac{\left(\frac{1-\rho_\pi^n}{1-\rho_\pi}\right)^2(1-\lambda)^2\sigma^2_{x\pi}}{(1-\lambda^2\rho_\pi^2)B^{2,\$}_{c,n}\sigma^2_{xc} + B^{2,\$}_{\pi,n}(1-\lambda)^2\sigma^2_{x\pi}}. \tag{35}$$

This expression suggests a key mechanism: when investors have sticky inflation beliefs (larger λ), the inflation variance ratio is attenuated. The intuition is that, with sticky beliefs, inflation news is not immediately incorporated into bond yields. In the limit, as λ approaches one, inflation news has vanishing impact on yields, and yield innovations are driven entirely by real rate news. Conversely, when $\lambda = 0$ (rational expectations), the full inflation signal is immediately incorporated into yields. This mechanism helps reconcile the empirical fact that innovations to inflation expectations account for only a small fraction of yield innovations.

5.7 Simulations

To illustrate the underlying mechanisms, I simulate a positive inflation shock in the model economy. For the sticky expectations model, I set the quarterly belief stickiness parameter

to $\lambda = \left(\frac{1.57}{1+1.57}\right)^{0.25} \approx 0.9$ based on the estimates in column (3) of Table 2, close to the value reported in Coibion and Gorodnichenko (2015). To maintain comparability with the literature, I adopt the remaining parameters from Bansal and Shaliastovich (2013), which are displayed in Table 10.

To provide a visual counterpart to the regression results in Table 3, Figure 5 plots the impulse response of the log one-period excess return on a 5-year bond to a one-standard-deviation expected inflation shock ($\sigma_{x\pi}$). When investors hold rational expectations, bond excess returns drop immediately following a positive expected inflation shock and then quickly revert to their steady-state values. In contrast, when investors hold sticky inflation expectations, the initial bond price decline is more muted, followed by a gradual decline over subsequent quarters as investors gradually realize their mistakes. As a result, an econometrician would observe that bond returns remain depressed for several quarters following inflation news, generating a negative relation between inflation news and future bond excess returns.

Moreover, the bond return predictability originates from the insufficient incorporation of inflation news into yield innovations. This relates to the critique of Duffee (2018), who shows that news about expected inflation accounts for only 10-20% of the variance of Treasury yield shocks, a finding at odds with standard macro-finance models. In particular, standard term structure models with recursive preferences and long-run risks counterfactually imply an inflation variance ratio that is too high, often exceeding one. The proposed sticky inflation belief hypothesis partially addresses this critique. To see this, Figure 6 plots the inflation variance ratio across bond maturities. The top panel plots the standard deviations of inflation news and yield innovations for the FIRE and sticky belief models. At all maturities, the magnitude of news about expected average inflation is smaller under sticky inflation expectations than that under FIRE. The bottom panel plots the model-implied inflation variance ratio, defined as the ratio of the variance of inflation news to the variance of yield innovations. As shown, the inflation variance ratio is lower when investors have sticky inflation expectations (dashed line) than under FIRE (solid line). Quantitatively, sticky inflation expectations bring the inflation variance ratio at the short end of the yield curve close to the empirical estimate of approximately 20% documented by Duffee (2018) but leave scope for improvement at the long end.[11] Thus, I propose sticky inflation beliefs as a complementary force contributing to the low inflation variance ratio, in addition to rational

[11] One potential extension that could bring the model-implied inflation variance ratio at longer maturities closer to its empirical counterpart is to incorporate overreaction in expectations of long-term interest rates, as in Wang (2021).

explanations (Gómez-Cram and Yaron, 2021; Schneider, 2022).

6 Conclusions

This study identifies a negative relation between inflation news and future bond excess returns, which is difficult to reconcile with rational explanations. The findings are more consistent with investors' inflation beliefs being sticky. Indeed, inflation forecast revisions positively predict subsequent forecast errors—a pattern inconsistent with FIRE but consistent with models of sticky information or noisy signals. The yield responses to inflation announcements confirm that bond yields do not fully adjust to inflation news on impact, with yield spreads continuing to drift in the direction of the surprise over subsequent trading days.

Sticky inflation beliefs reconcile the finding that bond excess returns are lower following positive inflation news. Moreover, they suggest that shocks to expected future inflation account for a smaller fraction of news about nominal yields. To provide a theoretical foundation for these findings, I develop a consumption-based asset pricing model with sticky beliefs about inflation. The model generates predictions that match the empirical patterns: (i) inflation forecast revisions predict realized forecast errors, (ii) inflation news negatively predicts future bond excess returns, (iii) yields positively predict future bond excess returns, and (iv) inflation variance ratios are attenuated. Overall, this study provides a parsimonious framework to understand how frictions in inflation expectations affect bond pricing and risk premia.

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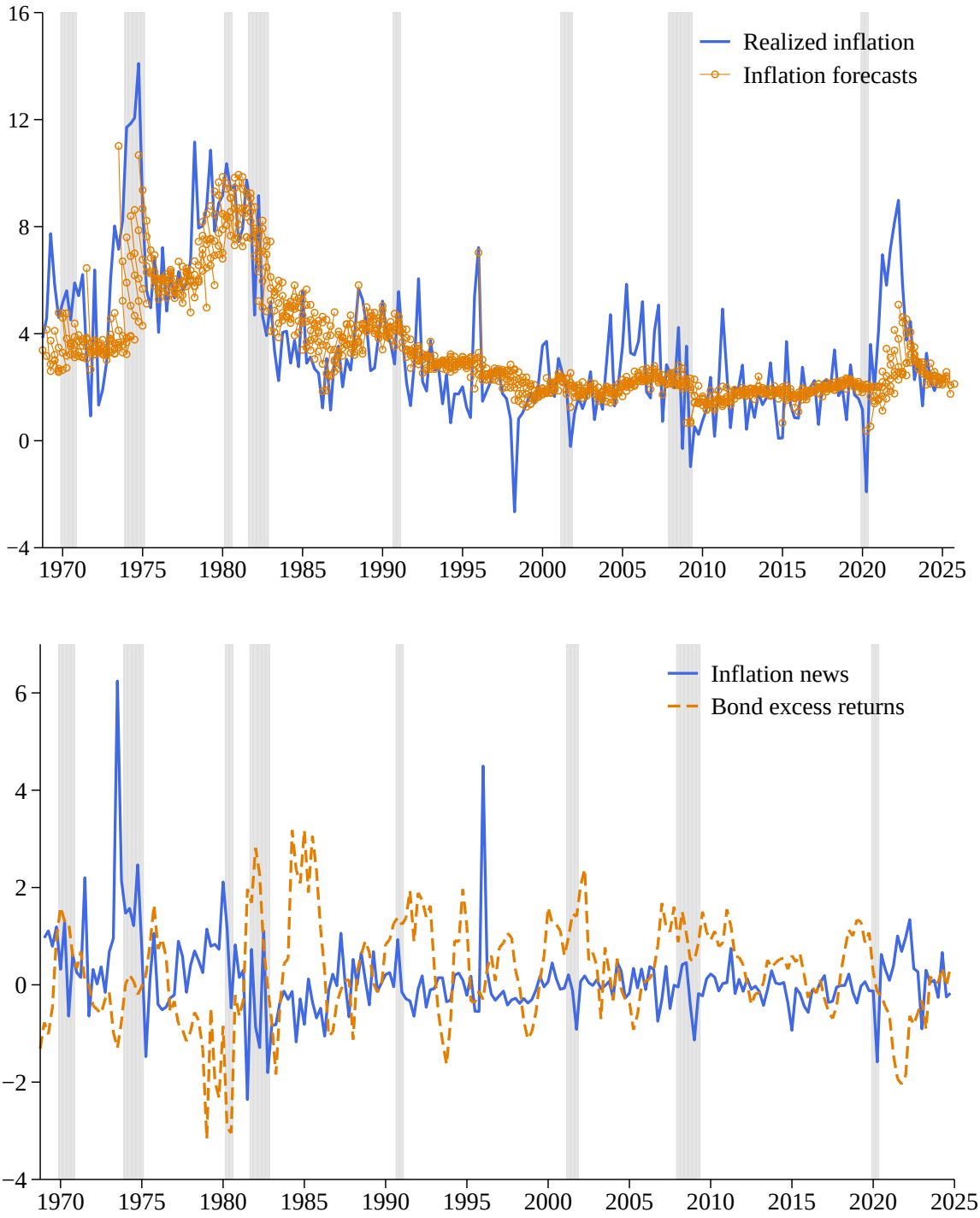


Figure 1: Inflation forecasts and annual bond excess returns. The top panel plots the current-quarter inflation and forecasts of inflation in the SPF. The quarter-over-quarter inflation forecasts (shown as circles) are for the current quarter through four quarters ahead. The bottom panel plots inflation news (forecast revisions for current-quarter inflation) and duration-standardized one-year excess holding-period returns on the 5-year bond. Data are quarterly observations from 1968Q4 to 2024Q4, expressed as annualized percentages. The shaded areas indicate NBER recessions.

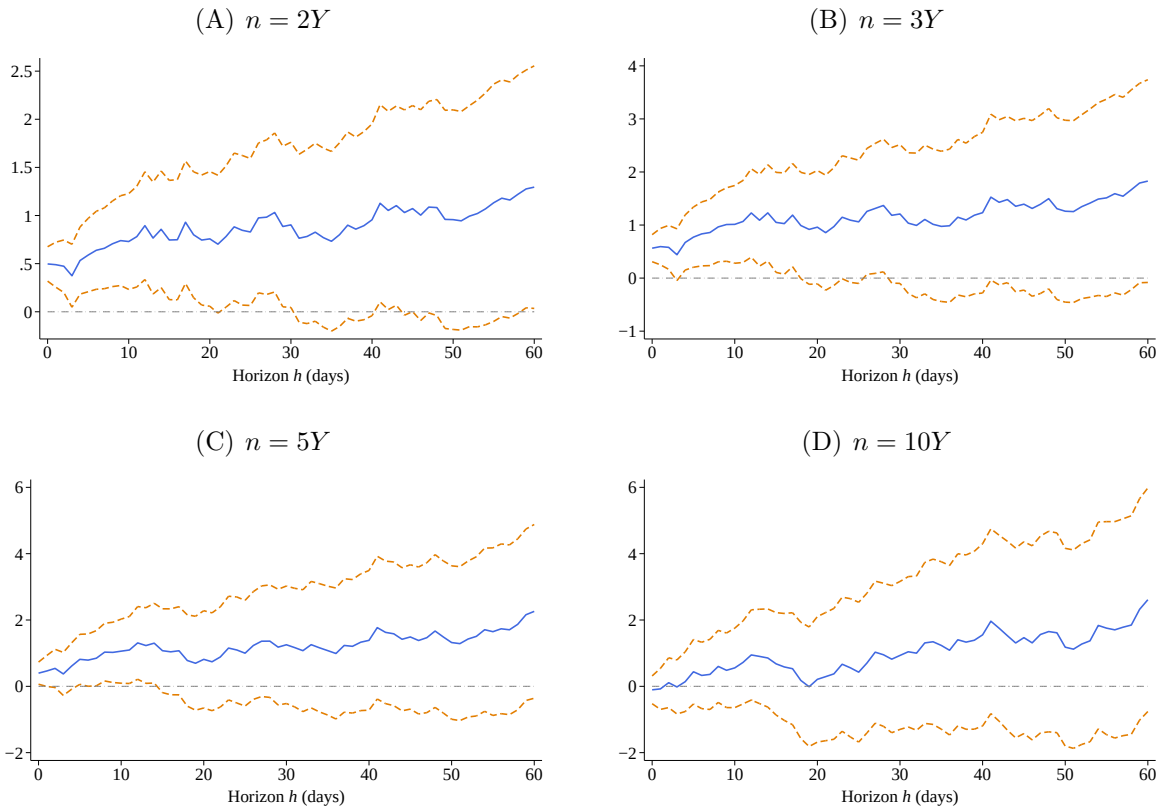


Figure 2: Inflation surprises and Treasury yield spreads. This figure plots the estimated effects (bps) of core CPI surprises on Treasury yield spreads for horizons h ranging from 0 to 60 business days. The estimated coefficients γ_h (solid blue line) and 90% confidence intervals (dashed orange lines) are obtained from the regression $\tilde{y}^{\$}_{\tau-1+h,n} = \alpha_h + \beta_h \tilde{y}^{\$}_{\tau-1,n} + \gamma_h \tilde{S}_\tau + \epsilon^h_\tau$ with restrictions $\alpha_h = 0$ and $\beta_h = 1$, where $\tilde{y}^{\$}_{\tau-1+h,n}$ denotes the yield spread between n-year and 1-year bonds, and $\tilde{S}_\tau$ is the realized month-over-month growth in core CPI minus the median Money Market Services (MMS) forecast. The estimated coefficient γ_h is interpreted as basis points per one-standard-deviation increase in the core CPI surprise. Standard errors are heteroskedasticity- and autocorrelation-consistent. The sample period spans from 1989 to 2025.

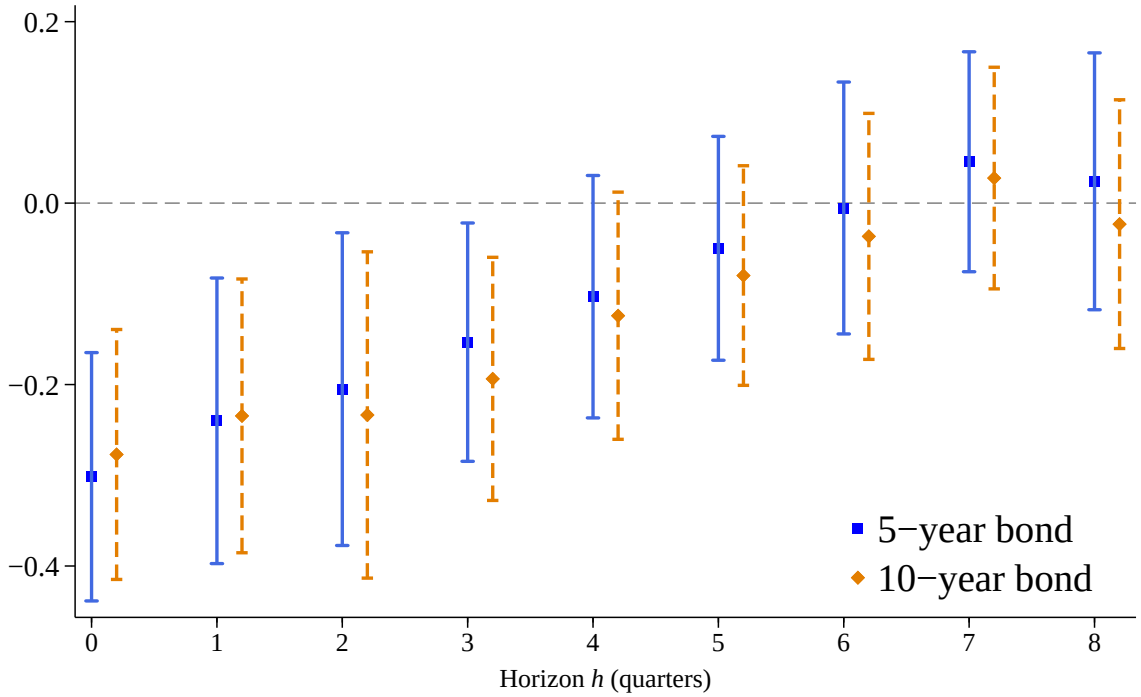


Figure 3: Inflation news and annual bond excess returns. The figure plots the coefficients β^h from regressing annual bond excess returns on inflation news: $rx^{\$}_{t+\frac{h}{4}\to t+1+\frac{h}{4},n}/n = \alpha + \beta^h(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \epsilon_{t+1+\frac{h}{4}}$, for horizons h from 0 to 8 quarters. The dependent variable is the annualized excess return (in percent) for an investor who purchases an n-year bond h quarters after the forecast revision and holds it for one year. The squares and diamonds represent estimates for the 5- and 10-year bonds, respectively. The independent variable is standardized to have zero mean and unit variance. The vertical bars indicate 90% confidence intervals based on heteroskedasticity- and autocorrelation-consistent standard errors with four lags.

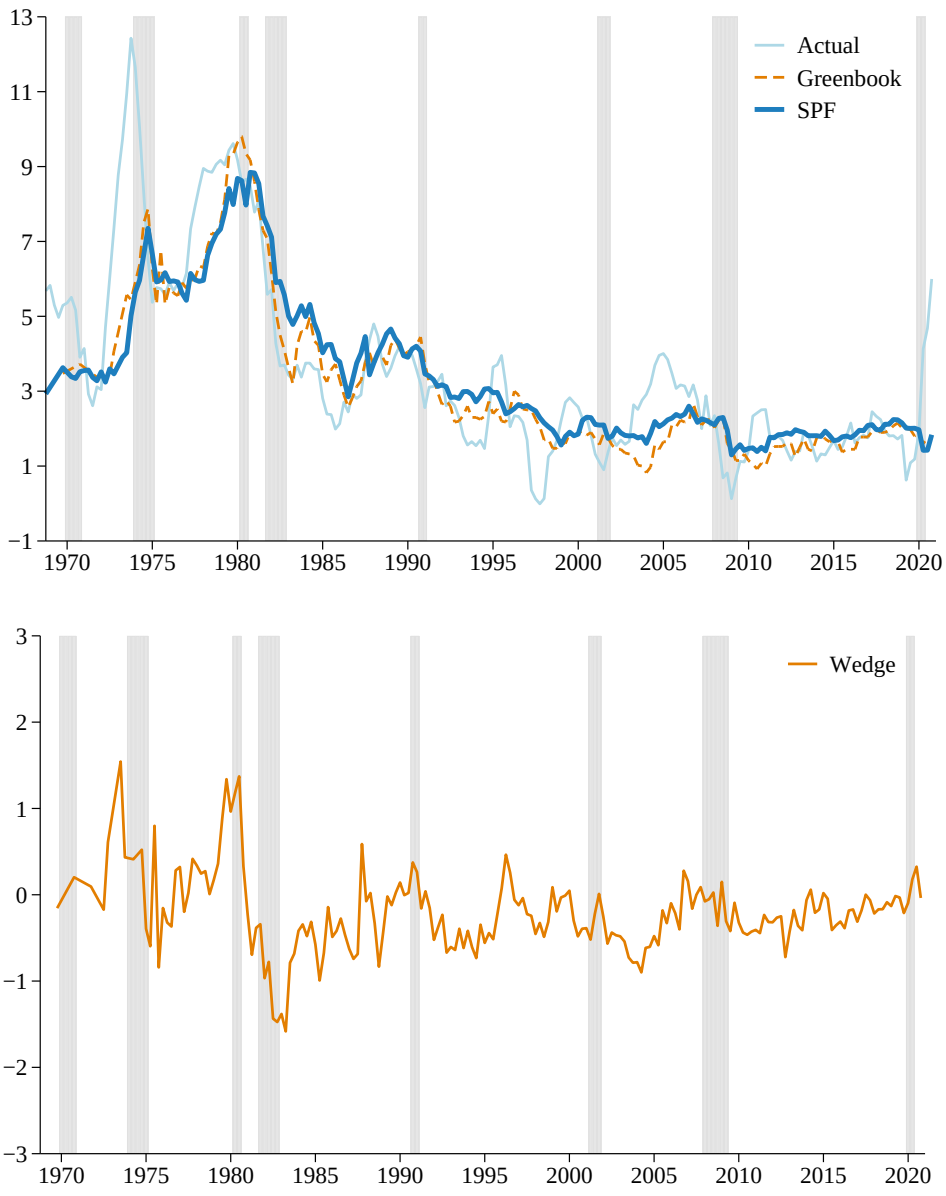


Figure 4: Inflation expectations: Greenbook versus SPF forecasts. The top panel plots the Greenbook forecasts, SPF forecasts, and corresponding actual values of one-year-ahead inflation. The scale is in annualized percent. The bottom panel shows the difference between the Greenbook and SPF forecasts (Equation (5)).

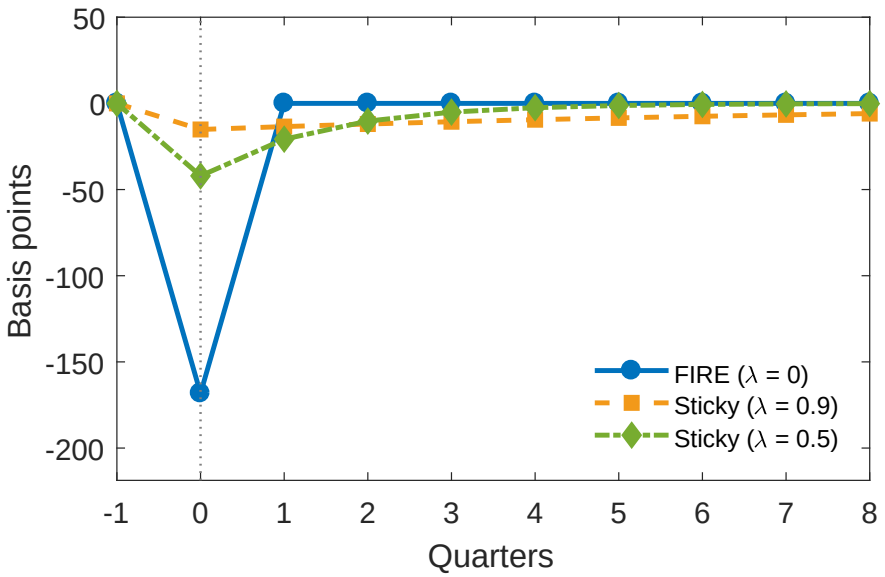


Figure 5: Impulse response to expected inflation shocks. The figure plots the impulse response of the log one-quarter excess return on 5-year bonds, measured as deviations from their steady-state values, to a one-standard-deviation expected inflation shock ($\sigma_{x\pi}$). The solid line corresponds to full-information rational expectations ($\lambda = 0$), the dashed line corresponds to sticky subjective inflation expectations ($\lambda = 0.9$), and the dash-dotted line corresponds to an intermediate case ($\lambda = 0.5$).

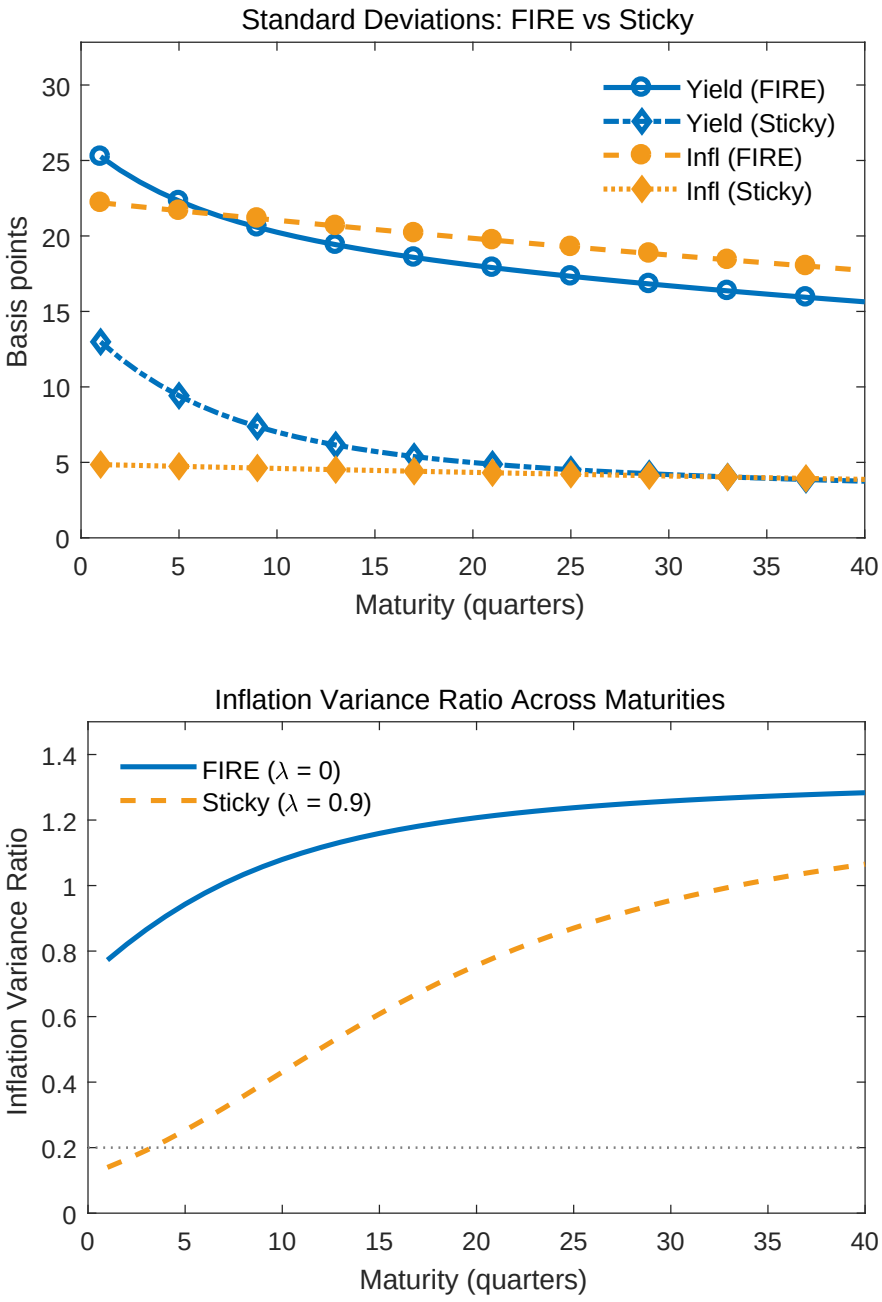


Figure 6: Standard deviations of inflation news and yield innovations. The top panel plots the standard deviations of inflation news and yield innovations across the models. Standard deviations are expressed in annualized basis points (quarterly data). The lines marked by the circles correspond to the FIRE model ($\lambda = 0$), while the lines marked by the diamonds correspond to the sticky inflation beliefs model ($\lambda = 0.9$). The bottom panel plots the model-implied inflation variance ratio, defined in Equation (32) as the variance of inflation news relative to that of yield innovations. The solid and the dashed lines correspond to the FIRE ($\lambda = 0$) and sticky inflation beliefs ($\lambda = 0.9$) models, respectively.

Table 1: Summary Statistics

The table reports summary statistics for bond excess returns. Panel A reports the mean, standard deviation, and autocorrelation for annual excess returns on n-year Treasury bonds, $rx^{\$}_{t\to t+1,n}$. Panel B reports the mean, standard deviation, and autocorrelation for duration-standardized excess returns, $rx^{\$}_{t\to t+1,n}/n$. AC(1) denotes the first-order autocorrelation coefficient (quarterly frequency and annual holding period). Panel C reports correlations between excess returns across maturities. Returns are expressed in percentages. The data are quarterly and cover the sample period from 1968Q4 to 2024Q4 (yield data from 2025 are used to calculate one-year holding-period returns).

Panel A: Bond excess returns					
	$rx^{\$}_{t\to t+1,2}$	$rx^{\$}_{t\to t+1,3}$	$rx^{\$}_{t\to t+1,4}$	$rx^{\$}_{t\to t+1,5}$	$rx^{\$}_{t\to t+1,10}$
Mean	0.44	0.78	1.08	1.33	2.11
Std. Dev.	1.72	3.10	4.27	5.34	10.22
AC(1)	0.77	0.77	0.77	0.77	0.77
Panel B: Duration-standardized excess returns					
Mean	0.22	0.26	0.27	0.27	0.21
Std. Dev.	0.86	1.03	1.07	1.07	1.02
AC(1)	0.77	0.77	0.77	0.77	0.77
Panel C: Correlation of excess returns					
$rx^{\$}_{t\to t+1,2}$	1.00				
$rx^{\$}_{t\to t+1,3}$	0.99	1.00			
$rx^{\$}_{t\to t+1,4}$	0.96	0.99	1.00		
$rx^{\$}_{t\to t+1,5}$	0.94	0.98	1.00	1.00	
$rx^{\$}_{t\to t+1,10}$	0.84	0.90	0.93	0.96	1.00

Table 2: FIRE Tests of Survey Expectations of Inflation

The table reports the coefficients in the following regressions ($h = 1, 2, 3$):

$$\pi_{t+\frac{h}{4}} - \tilde{E}_t(\pi_{t+\frac{h}{4}}) = \alpha + \beta \left(\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}}) \right) + \epsilon_{t,t+\frac{h}{4}},$$

$$\tilde{E}_{t+\frac{1}{4}}(\pi_{t+\frac{h}{4}}) - \tilde{E}_t(\pi_{t+\frac{h}{4}}) = \alpha + \beta \left(\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}}) \right) + \epsilon_{t,t+\frac{h}{4}},$$

where $\tilde{E}_t(\pi_{t+\frac{h}{4}})$ is the forecast of the cumulative h-quarter-ahead inflation rate at quarter t; $\pi_{t+\frac{h}{4}} - \tilde{E}_t(\pi_{t+\frac{h}{4}})$ is the forecast error for the h-quarter-ahead inflation rate, defined as realized inflation minus the forecast of the h-quarter-ahead inflation rate; and $\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}})$ is the forecast revision for the h-quarter-ahead inflation rate. Inflation forecasts are expressed in percent and are obtained from the SPF. The full sample covers 1969Q1 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Dependent Variables =	Forecast Errors			Forecast Revisions		
	h=1Q	h=2Q	h=3Q	h=1Q	h=2Q	h=3Q
	(1)	(2)	(3)	(4)	(5)	(6)
$\tilde{E}_t(\pi_{t+\frac{1}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{1}{4}})$	0.56			0.20		
	(2.16)			(2.95)		
$\tilde{E}_t(\pi_{t+\frac{2}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{2}{4}})$		1.06			0.32	
		(3.31)			(4.14)	
$\tilde{E}_t(\pi_{t+\frac{3}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{3}{4}})$			1.57			0.40
			(3.54)			(4.09)
Intercept	0.08	0.12	0.15	0.02	0.02	0.01
	(1.34)	(1.33)	(1.17)	(1.61)	(0.94)	(0.62)
Obs	224	224	218	223	223	218
R^2	0.06	0.16	0.24	0.09	0.16	0.21

Table 3: Inflation News and Bond Excess Returns

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(\pi_t)$ is the time t survey forecast of the current-quarter inflation rate, and $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of current-quarter inflation rate. Panel A reports results of univariate regressions. Panel B reports results of multivariate regressions that include the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury Bill rate TBL_t, inflation disagreement $Inflation_vol_t$, and real GDP growth disagreement $Real_vol_t$ as control variables V_t. The independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the SPF. The full sample spans 1969Q1 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.23	-0.28	-0.30	-0.28
	(-3.29)	(-3.49)	(-3.61)	(-3.30)
Obs	224	224	224	214
R^2	0.07	0.07	0.08	0.07

Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.22	-0.27	-0.28	-0.24
	(-3.87)	(-3.92)	(-3.93)	(-3.70)
CP_t	0.12	0.14	0.11	0.01
	(0.96)	(0.93)	(0.74)	(0.07)
CF_t	0.30	0.41	0.49	0.60
	(2.52)	(2.80)	(3.37)	(4.95)
τ_t^{CPI}	-0.43	-0.50	-0.44	-0.31
	(-2.32)	(-2.29)	(-2.03)	(-1.47)
GAP_t	-0.06	-0.03	0.04	0.10
	(-0.71)	(-0.29)	(0.34)	(1.02)
TBL_t	0.50	0.52	0.40	0.25
	(3.37)	(2.94)	(2.24)	(1.40)
$Inflation_vol_t$	0.05	0.05	0.04	0.04
	(0.75)	(0.73)	(0.54)	(0.50)
$Real_vol_t$	0.02	0.03	0.05	0.04
	(0.39)	(0.57)	(0.78)	(0.82)
Obs	224	224	224	214
R^2	0.38	0.37	0.38	0.42

Table 4: Evidence from the Michigan Survey of Consumers

The table reports the coefficients from the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n \, Household_news_t + \gamma^n V_t + \epsilon^n_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, and $Household_news_t$ is net unfavorable inflation news from the Michigan Survey of Consumers: the percentage of respondents reporting unfavorable price news minus the percentage reporting favorable price news. Panel A reports results of univariate regressions. Panel B reports results of multivariate regressions that include the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury Bill rate TBL_t, inflation disagreement $Inflation_vol_t$, and real GDP growth disagreement $Real_vol_t$ as control variables V_t. The independent variables are standardized to have zero mean and unit variance. The full sample spans 1978Q1 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n Household_news_t + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$Household_news_t$	-0.26	-0.30	-0.31	-0.31
	(-2.66)	(-2.54)	(-2.49)	(-2.72)
Obs	188	188	188	188
R^2	0.09	0.08	0.08	0.09
Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n Household_news_t + \gamma^n V_t + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$Household_news_t$	-0.25	-0.29	-0.28	-0.29
	(-2.54)	(-2.54)	(-2.63)	(-3.14)
CP_t	0.14	0.16	0.12	0.05
	(1.01)	(0.99)	(0.76)	(0.31)
CF_t	0.29	0.41	0.51	0.58
	(2.31)	(2.69)	(3.44)	(4.57)
τ_t^{CPI}	-0.67	-0.81	-0.79	-0.63
	(-3.04)	(-3.10)	(-3.01)	(-2.52)
GAP_t	0.01	0.04	0.10	0.19
	(0.15)	(0.36)	(0.82)	(1.67)
TBL_t	0.64	0.71	0.62	0.43
	(3.65)	(3.39)	(2.95)	(2.22)
$Inflation_vol_t$	0.20	0.25	0.26	0.27
	(2.20)	(2.41)	(2.46)	(2.65)
$Real_vol_t$	0.04	0.05	0.05	0.05
	(0.82)	(0.82)	(0.89)	(0.96)
Obs	188	188	188	188
R^2	0.38	0.38	0.40	0.44

Table 5: Inflation Belief Wedges and Bond Returns

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n \Delta_t(\pi_{t+1}) + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, and $\Delta_t(\pi_{t+1})$ is the difference between the Greenbook and SPF one-year-ahead inflation forecasts. Panel A reports results of univariate regressions. Panel B reports results of multivariate regressions that include the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury Bill rate TBL_t, inflation disagreement $Inflation_vol_t$, and real GDP growth disagreement $Real_vol_t$ as control variables V_t. The independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the Greenbook and the SPF. The full sample spans 1969Q4 to 2020Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n \Delta_t(\pi_{t+1}) + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\Delta_t(\pi_{t+1})$	-0.28	-0.32	-0.34	-0.34
	(-2.42)	(-2.19)	(-2.06)	(-2.09)
Obs	193	193	193	191
R^2	0.11	0.10	0.10	0.11
Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n \Delta_t(\pi_{t+1}) + \gamma^n V_t + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\Delta_t(\pi_{t+1})$	-0.24	-0.25	-0.23	-0.20
	(-2.20)	(-1.92)	(-1.71)	(-1.58)
CP_t	0.10	0.11	0.07	0.01
	(0.66)	(0.64)	(0.41)	(0.10)
CF_t	0.30	0.43	0.56	0.63
	(2.15)	(2.62)	(3.49)	(4.69)
τ_t^{CPI}	-0.45	-0.51	-0.44	-0.28
	(-2.23)	(-2.11)	(-1.83)	(-1.28)
GAP_t	-0.04	-0.01	0.05	0.13
	(-0.51)	(-0.13)	(0.46)	(1.34)
TBL_t	0.48	0.48	0.35	0.16
	(3.18)	(2.66)	(1.93)	(0.91)
$Inflation_vol_t$	0.10	0.12	0.12	0.11
	(1.44)	(1.32)	(1.18)	(1.18)
$Real_vol_t$	0.09	0.12	0.13	0.11
	(1.67)	(1.84)	(1.95)	(1.86)
Obs	193	193	193	191
R^2	0.36	0.36	0.38	0.43

Table 6: Controlling for Short-Rate Forecast Biases

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n Wedge_TBL_t + \delta^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of the current-quarter inflation rate, and $Wedge_TBL_t$ is the short-rate expectations wedge following Cieslak (2018), defined as the difference between an econometric forecast and the corresponding SPF forecast of the 3-month Treasury bill rate. The econometric forecast is obtained by regressing the one-year-ahead bill rate on lagged bill rates and employment growth over a 5-year rolling window. Panel A reports results from bivariate regressions of bond excess returns on inflation news and the short-rate expectations wedge. Panel B augments the specification with additional control variables V_t: the Cochrane and Piazzesi (2005) yield factor CP_t, the Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, the Cooper and Priestley (2009) output gap factor GAP_t, the 3-month Treasury bill rate TBL_t, inflation forecast disagreement $Inflation_vol_t$, and real GDP growth forecast disagreement $Real_vol_t$. All independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the SPF. The sample spans 1981Q3 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n Wedge_TBL_t + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.26	-0.30	-0.28	-0.22
	(-2.06)	(-2.10)	(-2.01)	(-1.64)
$Wedge_TBL_t$	-0.05	-0.09	-0.12	-0.12
	(-1.14)	(-1.62)	(-2.13)	(-2.22)
Obs	174	174	174	174
R^2	0.06	0.06	0.05	0.04

Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n Wedge_TBL_t + \delta^n V_t + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.17	-0.21	-0.23	-0.20
	(-2.56)	(-2.78)	(-2.94)	(-2.57)
$Wedge_TBL_t$	0.06	0.03	-0.02	-0.02
	(0.76)	(0.25)	(-0.18)	(-0.14)
Controls	Yes	Yes	Yes	Yes
Obs	174	174	174	174
R^2	0.46	0.42	0.39	0.38

Table 7: Real Rate News and Bond Excess Returns

The table reports the coefficients β^n in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)) + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(r_t)$ is the time t survey forecast of the current-quarter real rate (3-month T-bill rate minus GDP deflator inflation), and $\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)$ is the forecast revision. Panel A reports results of univariate regressions. Panel B reports results of multivariate regressions that include the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury Bill rate TBL_t, inflation disagreement $Inflation_vol_t$, and real GDP growth disagreement $Real_vol_t$ as control variables V_t. The independent variables are standardized to have zero mean and unit variance. Survey forecasts are obtained from the SPF. The full sample spans 1981Q4 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)) + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)$	0.04	0.05	0.05	0.06
	(0.53)	(0.56)	(0.56)	(0.69)
Obs	173	173	173	173
R^2	0.00	0.00	0.00	0.00

Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)) + \gamma^n V_t + \epsilon_{t+1}$

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(r_t) - \tilde{E}_{t-\frac{1}{4}}(r_t)$	0.09	0.12	0.12	0.10
	(1.83)	(1.88)	(1.72)	(1.43)
CP_t	0.12	0.14	0.11	0.00
	(0.93)	(0.89)	(0.60)	(0.02)
CF_t	0.24	0.34	0.45	0.55
	(2.05)	(2.36)	(2.99)	(3.94)
τ_t^{CPI}	-0.47	-0.62	-0.60	-0.34
	(-2.15)	(-2.25)	(-1.97)	(-1.06)
GAP_t	-0.05	-0.04	0.02	0.11
	(-0.65)	(-0.37)	(0.14)	(0.94)
TBL_t	0.67	0.77	0.67	0.38
	(4.00)	(3.56)	(2.73)	(1.47)
$Inflation_vol_t$	0.10	0.14	0.15	0.10
	(0.73)	(0.93)	(0.94)	(0.66)
$Real_vol_t$	-0.00	0.00	0.01	0.02
	(-0.06)	(0.03)	(0.22)	(0.48)
Obs	173	173	173	173
R^2	0.43	0.39	0.36	0.36

Table 8: Out-of-Sample Tests

The table reports the out-of-sample predictability of inflation news for bond excess returns. The rows labeled "ENC-NEW" report the test statistic proposed by Clark and McCracken (2001). The ENC-NEW test statistic for maturity n is given by:

$$\text{ENC-NEW}^{(n)} = P\frac{\sum_{t=1}^{T}\left(u_{t+4}^{2,(n)} - u_{t+4}^{(n)}\varepsilon_{t+4}^{(n)}\right)}{\sum_{t=1}^{T}\varepsilon_{t+4}^{2,(n)}},$$

where P is the number of 1-step-ahead predictions and $\varepsilon_t^{(n)}$ and $u_t^{(n)}$ denote the prediction errors from the unrestricted and restricted models, respectively. The rows labeled "R^2_{oos}" report the out-of-sample R^2 proposed by Campbell and Thompson (2008), expressed in percentage points. This measure compares the mean-squared prediction error of the predictor-based model with that of the historical-average benchmark:

$$R_{oos}^{2,(n)} = 1 - \frac{\sum_{t=1}^{T-4}\left(rx_{t\to t+4,n}^{\$} - \widehat{rx}_{\pi,t\to t+4,n}^{\$}\right)^2}{\sum_{t=1}^{T-4}\left(rx_{t\to t+4,n}^{\$} - \overline{rx}_{t\to t+4,n}^{\$}\right)^2},$$

where the time index t and sample size T are expressed in quarters. $\widehat{rx}_{\pi,t\to t+4,n}^{\$}$ is the forecast of the annual excess return based on time-t inflation news, where the parameters of the predictive model are estimated using inflation news up to time $t-4$ and returns realized up to time t. $\overline{rx}_{t\to t+4,n}^{\$}$ is the return forecast using the historical average excess return estimated through time t. The rows labeled "MSE-F" report the test statistic of McCracken (2007), and the rows labeled "p-value" report p-values computed following Clark and West (2007). All out-of-sample forecasts are formed recursively.

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
Panel A: Out-of-sample period: 1978Q1-2024Q4				
ENC-NEW	9.32	9.86	10.44	8.71
MSE-F	10.57	11.62	12.33	9.65
R^2_{oos}	5.32	5.82	6.15	4.88
p-value	0.00	0.00	0.00	0.00
Panel B: Out-of-sample period: 1985Q1-2024Q4				
ENC-NEW	8.65	9.27	9.73	8.21
MSE-F	6.53	8.51	9.74	7.56
R^2_{oos}	3.92	5.05	5.74	4.51
p-value	0.00	0.00	0.00	0.00

Table 9: Inflation News, Belief Stickiness, and Bond Excess Returns

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta_1^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \beta_2^n\hat{\lambda}_t + \beta_3^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) \times \hat{\lambda}_t + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond; $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of current-quarter inflation rate; and $\hat{\lambda}_t$ measures the degree of stickiness in inflation beliefs, converted from the slope coefficient of a 10-year rolling regression of 3-quarter-ahead inflation forecast errors on forecast revisions. The regressions include the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury Bill rate TBL_t, inflation disagreement $Inflation_vol_t$, and real GDP growth disagreement $Real_vol_t$ as control variables V_t. The independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the SPF. The sample spans 1978Q4 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.32	-0.39	-0.39	-0.33
	(-4.68)	(-4.60)	(-4.49)	(-4.01)
$\hat{\lambda}_t$	-0.04	-0.01	0.03	0.03
	(-0.44)	(-0.10)	(0.25)	(0.31)
$(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) \times \hat{\lambda}_t$	-0.14	-0.16	-0.15	-0.14
	(-3.79)	(-3.71)	(-3.58)	(-3.23)
Controls	Yes	Yes	Yes	Yes
Obs	185	185	185	185
R^2	0.41	0.41	0.43	0.45

Table 10: Model parameters

This table displays model parameter values for U.S. quarterly data used in this study based on Table 4 of Bansal and Shaliastovich (2013). The stickiness of inflation expectations (λ) is set to 0.9 in the case of sticky expectations.

Parameter	Symbol	Value
Panel A: Preferences		
Subjective discount factor	δ	0.994
Risk aversion	γ	20.90
Elasticity of intertemporal substitution	ψ	1.81
Panel B: Consumption		
Mean consumption growth	μ_c	4.9e-03
Volatility of consumption shocks	σ_c	4.6e-03
Persistence of expected consumption growth	ρ_c	0.810
Loading of expected growth on expected inflation	$\rho_{c\pi}$	-0.047
Volatility of expected consumption growth shocks	σ_{xc}	1.09e-03
Panel C: Inflation		
Mean inflation	μ_π	9.0e-03
Volatility of inflation shocks	σ_π	5.5e-03
Persistence of expected inflation	ρ_π	0.988
Volatility of expected inflation shocks	$\sigma_{x\pi}$	1.11e-03
Stickiness of inflation expectations	λ	0.9

Appendix for
"Imperfect Inflation Expectations and Treasury Bond Returns"

Appendices A and B provide additional results and model derivations, respectively.

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A Additional Results

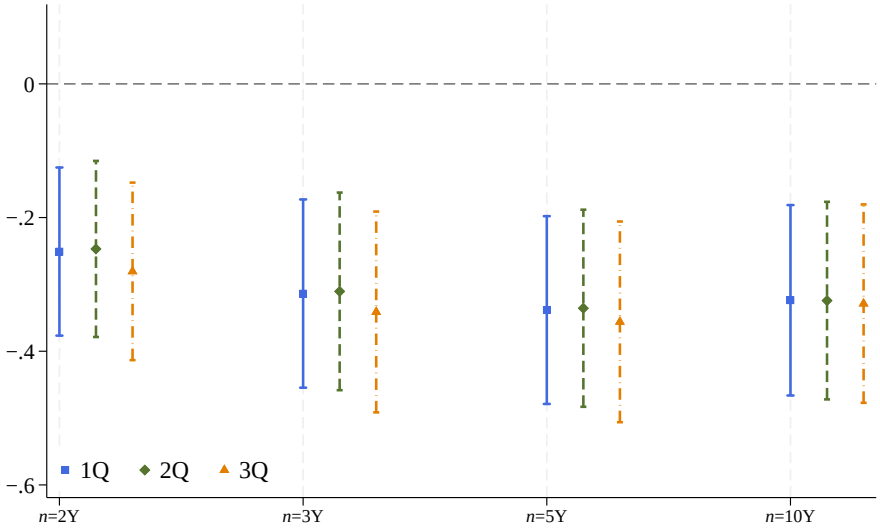


Figure A1: Inflation news across horizons and bond returns. The figure plots the coefficients β^n from regressions of annual bond excess returns on h-quarter-ahead inflation forecast revisions, $rx^{\$}_{t\to t+1,n}/n = \alpha + \beta^n(\tilde{E}_t(\pi_{t+\frac{h}{4}}) - \tilde{E}_{t-\frac{1}{4}}(\pi_{t+\frac{h}{4}})) + \epsilon_{t+1}$ for $n = 2, 3, 5, 10$ years. Squares, diamonds, and triangles denote coefficients on 1-, 2-, and 3-quarter-ahead forecast revisions, respectively. Revisions are standardized to have zero mean and unit variance. The vertical bars are 90% confidence intervals computed using Newey-West standard errors with four lags.

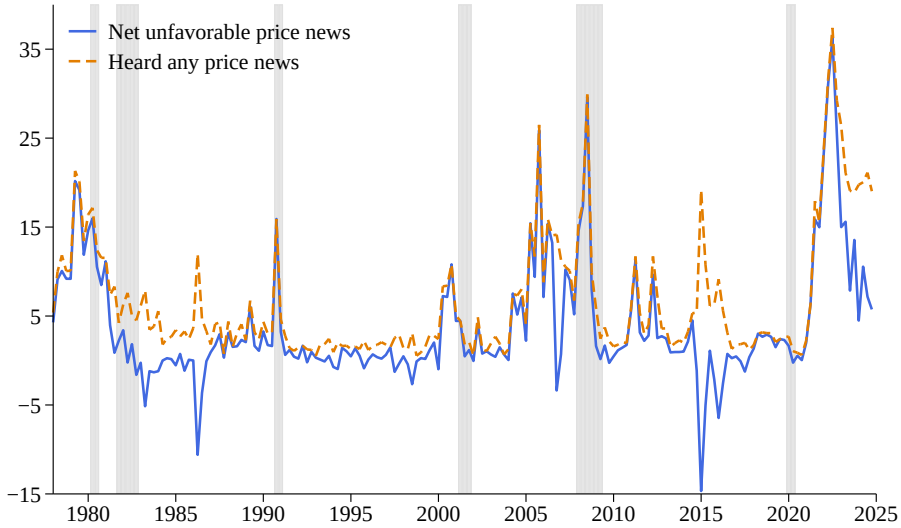


Figure A2: Household inflation news in the Michigan Survey of Consumers. The solid line plots net unfavorable price news: the percentage of respondents hearing unfavorable news about higher prices minus the percentage hearing favorable news about lower prices. The dashed line plots the percentage hearing any price news. Both series are weighted percentages of all respondents and count mentions in the survey's responses to the question, "During the last few months, have you heard of any favorable or unfavorable changes in business conditions? What did you hear?" Quarterly data are from the first-month survey of each quarter and span 1978Q1 to 2024Q4. Shaded areas indicate NBER recessions.

Table A1: Robustness: Inference with Overlapping Returns

The table reports coefficients from the regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the annual-holding-period excess return on an n-year bond and $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the revision of the SPF current-quarter inflation forecast. Panel A reports univariate regressions. Panel B adds as controls V_t the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap GAP_t, 3-month T-Bill rate TBL_t, and inflation and real GDP growth disagreement ($Inflation_vol_t$, $Real_vol_t$). Independent variables are standardized. The sample spans 1969Q1 to 2024Q4. The t-statistics in parentheses are based on the reverse regression delta method of Wei and Wright (2013).

Panel A: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.23	-0.28	-0.30	-0.28
	(-2.22)	(-2.46)	(-2.64)	(-2.71)
Obs	224	224	224	214
R^2	0.07	0.07	0.08	0.07
Panel B: $rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1}$				
	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.22	-0.27	-0.28	-0.24
	(-2.71)	(-2.88)	(-3.01)	(-2.95)
CP_t	0.12	0.14	0.11	0.01
	(-0.18)	(-0.03)	(0.04)	(-0.10)
CF_t	0.30	0.41	0.49	0.60
	(2.76)	(3.09)	(3.60)	(4.26)
τ_t^{CPI}	-0.43	-0.50	-0.44	-0.31
	(-1.57)	(-1.56)	(-1.38)	(-1.04)
GAP_t	-0.06	-0.03	0.04	0.10
	(-0.30)	(-0.03)	(0.41)	(0.94)
TBL_t	0.50	0.52	0.40	0.25
	(3.23)	(2.86)	(2.19)	(1.23)
$Inflation_vol_t$	0.05	0.05	0.04	0.04
	(0.37)	(0.40)	(0.40)	(0.42)
$Real_vol_t$	0.02	0.03	0.05	0.04
	(0.49)	(0.66)	(0.76)	(0.63)
Obs	224	224	224	214
R^2	0.38	0.37	0.38	0.42

Table A2: Inflation News, Forecast Errors, and Bond Excess Returns

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n(\pi_{t+1} - \tilde{E}_t(\pi_{t+1})) + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of current-quarter inflation rate, and $\pi_{t+1} - \tilde{E}_t(\pi_{t+1})$ is the error in SPF one-year-ahead inflation forecasts. The independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the SPF. The full sample spans 1969Q1 to 2024Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.08	-0.11	-0.13	-0.13
	(-1.49)	(-1.65)	(-1.80)	(-1.80)
$\pi_{t+1} - \tilde{E}_t(\pi_{t+1})$	-0.37	-0.42	-0.41	-0.36
	(-4.11)	(-4.04)	(-4.12)	(-4.13)
Obs	215	215	215	209
R^2	0.22	0.21	0.20	0.18

Table A3: Inflation News and Bond Excess Returns: Additional Factors

The table reports the coefficients in the following regressions ($n = 2, 3, 5, 10$):

$$rx^{\$}_{t\to t+1,n}/n = \alpha^n + \beta^n(\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)) + \gamma^n V_t + \epsilon_{t+1},$$

where $rx^{\$}_{t\to t+1,n}/n$ is the excess return for an annual holding period on an n-year bond, $\tilde{E}_t(\pi_t)$ is the time t survey forecast of the current-quarter inflation rate, and $\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$ is the forecast revision of the current-quarter inflation rate. V_t is the vector of control variables: the Cochrane and Piazzesi (2005) yield factor CP_t, Cieslak and Povala (2015) cycle factor CF_t and trend inflation τ_t^{CPI}, Cooper and Priestley (2009) output gap factor GAP_t, 3-month Treasury bill rate TBL_t, inflation disagreement $Inflation_vol_t$, real GDP growth disagreement $Real_vol_t$, the modified Ludvigson and Ng (2011) factor LN_t, the hidden component H_t of the Duffee (2011) one-dimensional term premium factor, and the Berardi et al. (2021) convergence gap CG_t. The independent variables are standardized to have zero mean and unit variance. Inflation forecasts are obtained from the SPF. The full sample spans 1969Q1 to 2013Q4. Newey-West four-lag adjusted t-statistics are in parentheses.

	n=2Y	n=3Y	n=5Y	n=10Y
	(1)	(2)	(3)	(4)
$\tilde{E}_t(\pi_t) - \tilde{E}_{t-\frac{1}{4}}(\pi_t)$	-0.14	-0.17	-0.17	-0.15
	(-3.47)	(-3.54)	(-3.57)	(-3.52)
CP_t	-0.15	-0.23	-0.29	-0.33
	(-0.80)	(-1.11)	(-1.52)	(-1.94)
CF_t	0.48	0.67	0.81	0.91
	(3.20)	(3.91)	(5.09)	(7.06)
τ_t^{CPI}	0.01	0.11	0.23	0.32
	(0.04)	(0.33)	(0.77)	(1.11)
GAP_t	0.03	0.09	0.17	0.24
	(0.30)	(0.80)	(1.54)	(2.19)
TBL_t	0.14	0.11	0.04	-0.05
	(0.41)	(0.29)	(0.12)	(-0.17)
$Inflation_vol_t$	0.13	0.15	0.14	0.16
	(1.90)	(1.95)	(1.82)	(1.99)
$Real_vol_t$	0.36	0.67	0.94	0.82
	(0.90)	(1.43)	(1.98)	(1.56)
LN_t	0.21	0.27	0.30	0.31
	(2.56)	(2.77)	(2.89)	(2.61)
H_t	-0.16	-0.21	-0.21	-0.16
	(-1.09)	(-1.25)	(-1.26)	(-0.95)
CG_t	-0.20	-0.20	-0.16	-0.09
	(-1.54)	(-1.32)	(-0.95)	(-0.54)
Obs	180	180	180	170
R^2	0.47	0.49	0.52	0.57

Table A4: Comparison of Inflation Expectations: Greenbook versus SPF

The table reports coefficients from the following regression specifications:

$$\pi_{t+1} = \alpha + \beta E_t(\pi_{t+1}) + \epsilon_{t+1},$$
$$\pi_{t+1} = \alpha + \beta \tilde{E}_t(\pi_{t+1}) + \epsilon_{t+1},$$
$$\pi_{t+1} = \alpha + \beta_1 E_t(\pi_{t+1}) + \beta_2 \tilde{E}_t(\pi_{t+1}) + \epsilon_{t+1},$$
$$\pi_{t+1} - \tilde{E}_t(\pi_{t+1}) = \alpha + \beta \Delta_t(\pi_{t+1}) + \epsilon_{t+1},$$

where π_{t+1} is the realized 1-year-ahead inflation rate, $E_t(\pi_{t+1})$ is the Greenbook forecast of the 1-year-ahead inflation rate, $\tilde{E}_t(\pi_{t+1})$ is the SPF forecast of the 1-year-ahead inflation rate, and $\Delta_t(\pi_{t+1}) = E_t(\pi_{t+1}) - \tilde{E}_t(\pi_{t+1})$ is the wedge between the Greenbook and SPF forecasts. Columns (1)-(3) report results with realized inflation as the dependent variable; column (4) reports results with the SPF forecast error as the dependent variable. Inflation forecasts are expressed in percent. The sample spans 1969Q4 to 2020Q4 and includes only quarters for which both forecasts are available. Newey-West four-lag adjusted t-statistics are in parentheses.

Dependent Variable =	Realized Inflation π_{t+1}			SPF Forecast Error $\pi_{t+1} - \tilde{E}_t(\pi_{t+1})$
	(1)	(2)	(3)	(4)
$E_t(\pi_{t+1})$	0.99		1.52	
	(12.47)		(3.63)	
$\tilde{E}_t(\pi_{t+1})$		1.02	-0.58	
		(10.90)	(-1.42)	
$\Delta_t(\pi_{t+1})$				1.45
				(3.21)
Obs	193	193	193	193
R^2	0.73	0.67	0.74	0.22

B Model Solution

Appendix B provides the detailed solution to the consumption-based asset pricing model described in the main paper.

B.1 Model Setup

The assumed stochastic processes for the logarithm of consumption growth, Δc_{t+1}, and logarithm of the inflation rate, π_{t+1}, are

$$\begin{aligned}\Delta c_{t+1} &= \mu_c + x_{c,t} + \sigma_c \eta_{c,t+1},\\ \pi_{t+1} &= \mu_\pi + x_{\pi,t} + \sigma_\pi \eta_{\pi,t+1},\end{aligned}$$

where the predictable components evolve as follows:

$$\begin{aligned}x_{c,t+1} &= \rho_c x_{c,t} + \rho_{c\pi}\tilde{x}_{\pi,t} + \sigma_{xc}\eta_{xc,t+1}\\ x_{\pi,t+1} &= \rho_\pi x_{\pi,t} + \sigma_{x\pi}\eta_{x\pi,t+1}.\end{aligned} \tag{A1}$$

Here, $\tilde{x}_{\pi,t}$ captures the representative agent's subjective inflation beliefs, which influence household consumption decisions via inflation non-neutrality ($\rho_{c\pi}$). Specifically, the representative agent is assumed to hold sticky inflation beliefs that follow:

$$\begin{aligned}\tilde{E}_t[x_{\pi,t+1}] &= (1-\lambda)E_t[x_{\pi,t+1}] + \lambda\tilde{E}_{t-1}[x_{\pi,t+1}]\\ &= \rho_\pi \tilde{x}_{\pi,t}, \qquad \text{(A2)}\\ \tilde{x}_{\pi,t} &\equiv (1-\lambda)\sum_{j=0}^{\infty}(\lambda\rho_\pi)^j x_{\pi,t-j}. \qquad \text{(A3)}\end{aligned}$$

A tilde on the expectation operator indicates that the expectation is taken under the subjective measure. Note that inflation expectations are sticky relative to the true dynamics when $\lambda > 0$. When $\lambda = 0$, the inflation belief is rational and the model degenerates to the constant volatility setting of Bansal and Shaliastovich (2013).

Under the representative agent's beliefs, the shock to expected inflation is

$$\begin{aligned}\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1} &\equiv x_{\pi,t+1} - \tilde{E}_t[x_{\pi,t+1}]\\ &= \rho_\pi(x_{\pi,t} - \tilde{x}_{\pi,t}) + \sigma_{x\pi}\eta_{x\pi,t+1},\end{aligned} \tag{A4}$$

where $\rho_\pi(x_{\pi,t} - \tilde{x}_{\pi,t})$ is the representative agent's forecast bias. Therefore, the dynamics of expected

inflation under the representative agent's beliefs can be rewritten as

$$\tilde{x}_{\pi,t+1} = \rho_\pi \tilde{x}_{\pi,t} + (1-\lambda)\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1}. \tag{A5}$$

Given the assumed preferences for the representative agent, the logarithm of the real SDF is given by

$$m_{t+1} = \theta \ln \delta - \frac{\theta}{\psi}\Delta c_{t+1} + (\theta - 1) r_{c,t+1}, \tag{A6}$$

with $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$.

To obtain an analytical model solution, I rely on an approximation proposed by Campbell and Shiller (1988), which involves a log-linear Taylor expansion of $r_{c,t+1}$:

$$r_{c,t+1} = \kappa_0 + \kappa_1 pc_{t+1} - pc_t + \Delta c_{t+1}, \tag{A7}$$

where pc_{t+1} is the log price-to-consumption ratio and κ_0 and κ_1 are constants endogenously determined by the unconditional mean $\bar{pc}$ of the economy:

$$\kappa_1 = \frac{e^{\bar{pc}}}{1 + e^{\bar{pc}}}, \qquad \kappa_0 = \ln(1 + e^{\bar{pc}}) - \kappa_1 \bar{pc}. \tag{A8}$$

The Euler equation is

$$\tilde{E}_t[\exp(m_{t+1} + r_{t+1})] = 1, \tag{A9}$$

which is used to price any asset return r_{t+1} in the economy.

B.2 Solution for the Price-Consumption Ratio

In equilibrium, the price-consumption ratio is linear in the state variables:

$$pc_t = A_0 + A_c x_{c,t} + A_\pi \tilde{x}_{\pi,t}. \tag{A10}$$

To solve for the coefficients, I substitute the approximation (A7) and the SDF (A6) into the Euler equation (A9) and impose $\tilde{E}_t[\exp(m_{t+1} + r_{c,t+1})] = 1$. To satisfy the Euler equation at all times,

the following system of equations in A_c, A_π, and A_0 must hold:

$$(x_{ct}) : 0 = \left(1 - \frac{1}{\psi}\right) - A_c\left(1 - \kappa_1\rho_c\right) \tag{A11}$$

$$(\tilde{x}_{\pi t}) : 0 = \left(1 - \frac{1}{\psi}\right)\kappa_1\rho_{c\pi} - A_\pi(1 - \kappa_1\rho_\pi)(1 - \kappa_1\rho_c) \tag{A12}$$

$$\begin{aligned}(const) : 0 = \frac{\theta^2}{2}\left(1 - \frac{1}{\psi}\right)^2\sigma_c^2 - \frac{\theta}{\psi}\mu_c + \theta\left(\kappa_1 - 1\right)A_0 + \theta\left(\kappa_0 + \mu_c + \ln\delta\right) \\ + \frac{\theta^2}{2}\kappa_1^2A_c^2\sigma_{xc}^2 + \frac{\theta^2}{2}\kappa_1^2A_\pi^2(1-\lambda)^2\sigma_{x\pi}^2.\end{aligned} \tag{A13}$$

Using undetermined coefficients, I obtain the loadings and constant term

$$\begin{aligned}A_c &= \frac{1 - \frac{1}{\psi}}{1 - \kappa_1\rho_c}, \qquad A_\pi = \frac{\kappa_1\rho_{c\pi}(1 - \frac{1}{\psi})}{(1 - \kappa_1\rho_\pi)(1 - \kappa_1\rho_c)}, \\ A_0 &= \frac{\ln\delta + \kappa_0 + (1 - \frac{1}{\psi})\mu_c + \frac{\theta}{2}(1 - \frac{1}{\psi})^2\sigma_c^2 + \frac{\theta}{2}\kappa_1^2A_c^2\sigma_{xc}^2 + \frac{\theta}{2}\kappa_1^2A_\pi^2(1-\lambda)^2\sigma_{x\pi}^2}{1 - \kappa_1}.\end{aligned} \tag{A14}$$

B.3 Expression for the Real and Nominal SDF

Evaluating $r_{c,t+1}$ with the solution derived for pc_t and using the assumed process for consumption growth, the real SDF in Equation (A6) can be rewritten in terms of the state variables and underlying shocks (risk) under the agent's subjective expectations. Specifically, under the representative agent's beliefs, the real and nominal SDFs satisfy

$$\begin{aligned}m_{t+1} &= m_0 + m_c x_{c,t} + m_\pi\tilde{x}_{\pi,t} - \lambda_c\sigma_c\eta_{c,t+1} - \lambda_\pi\sigma_\pi\eta_{\pi,t+1} - \lambda_{xc}\sigma_{xc}\eta_{xc,t+1} - \lambda_{x\pi}\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1}, \\ m_{t+1}^\$ &= m_0^\$ + m_c^\$ x_{c,t} + m_\pi^\$\tilde{x}_{\pi,t} - \lambda_c^\$\sigma_c\eta_{c,t+1} - \lambda_\pi^\$\sigma_\pi\eta_{\pi,t+1} - \lambda_{xc}^\$\sigma_{xc}\eta_{xc,t+1} - \lambda_{x\pi}^\$\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1},\end{aligned} \tag{A15}$$

where the λ's represent the market price of each source of risk. The discount factor parameters and market prices of risks are equal to

$$\begin{aligned}&m_0 = \theta\ln\delta - \gamma\mu_c + (\theta-1)\kappa_0 + (\theta-1)(\kappa_1-1)A_0, \quad m_c = -\frac{1}{\psi}, \quad m_\pi = 0, \\ &\lambda_c = \gamma, \quad \lambda_\pi = 0, \quad \lambda_{xc} = (1-\theta)\kappa_1A_c, \quad \lambda_{x\pi} = (1-\theta)(1-\lambda)\kappa_1A_\pi\end{aligned} \tag{A16}$$

and the nominal ones are

$$\begin{aligned}&m_0^\$ = m_0 - \mu_\pi, \qquad m_c^\$ = -\frac{1}{\psi}, \quad m_\pi^\$ = m_\pi - 1, \qquad \lambda_c^\$ = \gamma, \\ &\lambda_\pi^\$ = \lambda_\pi + 1, \quad \lambda_{xc}^\$ = (1-\theta)\kappa_1A_c, \quad \lambda_{x\pi}^\$ = (1-\theta)(1-\lambda)\kappa_1A_\pi.\end{aligned} \tag{A17}$$

B.4 Solution for Equilibrium Nominal and Real Bond Yields

Next, I solve for the yields of nominal and real zero-coupon bonds of different maturities. Let $P^{\$}_{t,n}$ be the time-t price of an n-th period zero-coupon nominal bond that pays one unit of numeraire in n periods. Using the Euler equation, I can write the price in logs recursively as

$$p^{\$}_{t,n} = \ln \tilde{E}_t \left[\exp\left(m^{\$}_{t+1} + p^{\$}_{t+1,n-1} \right) \right], \tag{A18}$$

where I define $p^{\$}_{t,n} = \ln P^{\$}_{t,n}$. Conjecture that $p^{\$}_{t,n}$ is a linear function of the state variables:

$$p^{\$}_{t,n} = -B^{\$}_{0,n} - B^{\$}_{c,n} x_{c,t} - B^{\$}_{\pi,n} \tilde{x}_{\pi,t}. \tag{A19}$$

To solve for the B coefficients, I again use the method of undetermined coefficients and it follows that

$$\begin{aligned}
B^{\$}_{c,n} &= \rho_c B^{\$}_{c,n-1} - m^{\$}_c, \\
B^{\$}_{\pi,n} &= \rho_\pi B^{\$}_{\pi,n-1} + \rho_{c\pi} B^{\$}_{c,n-1} - m^{\$}_\pi, \\
B^{\$}_{0,n} &= B^{\$}_{0,n-1} - \frac{1}{2}[\lambda^{2,\$}_c \sigma^2_c + \lambda^{2,\$}_\pi \sigma^2_\pi + (\lambda^{\$}_{xc} + B^{\$}_{c,n-1})^2 \sigma^2_{xc} \\
&\quad + (\lambda^{\$}_{x\pi} + (1-\lambda) B^{\$}_{\pi,n-1})^2 \sigma^2_{x\pi}] - m^{\$}_0,
\end{aligned} \tag{A20}$$

with initial conditions $B^{\$}_{0,0} = 0$, $B^{\$}_{c,0} = 0$, and $B^{\$}_{\pi,0} = 0$. Iterating on Equation (A20), I obtain that

$$\begin{aligned}
B^{\$}_{c,n} &= \frac{1}{\psi} \frac{1-\rho^n_c}{1-\rho_c}, \\
B^{\$}_{\pi,n} &= \frac{1-\rho^n_\pi}{1-\rho_\pi} + \frac{\rho_{c\pi}}{\psi(\rho_c - \rho_\pi)} \left[\frac{\rho_c - \rho^n_c}{1-\rho_c} - \frac{\rho_\pi - \rho^n_\pi}{1-\rho_\pi} \right].
\end{aligned} \tag{A21}$$

Given the price of a zero-coupon bond, bond yields are defined as $y^{\$}_{t,n} = -\frac{1}{n} \ln P^{\$}_{t,n}$. Hence,

$$y^{\$}_{t,n} = \frac{1}{n}(B^{\$}_{0,n} + \underbrace{B^{\$}_{c,n}}_{>0} x_{c,t} + \underbrace{B^{\$}_{\pi,n}}_{>0} \tilde{x}_{\pi,t}). \tag{A22}$$

Similar equations are obtained for the real bond coefficients using the real discount factor parameters.

B.5 Expected Excess Bond Returns

Define the excess log return on buying an n-period bond at time t and selling it at time $t+1$ as an $n-1$ period bond:

$$rx^{\$}_{t \to t+1,n} = n y^{\$}_{t,n} - (n-1) y^{\$}_{t+1,n-1} - y^{\$}_{t,1}. \tag{A23}$$

As with the nominal SDF, the excess log return under the representative agent's subjective expectations can be expressed in terms of a time t expected component and return innovations:

$$rx^{\$}_{t\to t+1,n} - \tilde{E}_t rx^{\$}_{t\to t+1,n} = -B^{\$}_{c,n-1}\sigma_{xc}\eta_{xc,t+1} - (1-\lambda)B^{\$}_{\pi,n-1}\sigma_{x\pi}\tilde{\eta}_{x\pi,t+1}. \tag{A24}$$

In the model, one-period expected excess return on nominal bonds is determined by the negative covariation between innovations to excess log returns and the innovations to the nominal SDF in Equation (A15):

$$\begin{aligned}\tilde{E}_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}\tilde{Var}_t(rx^{\$}_{t\to t+1,n}) &= -\tilde{Cov}_t\left(m^{\$}_{t+1}, rx^{\$}_{t\to t+1,n}\right) \\ &= -\lambda^{\$}_{xc}B^{\$}_{c,n-1}\sigma^2_{xc} - (1-\lambda)\lambda^{\$}_{x\pi}B^{\$}_{\pi,n-1}\sigma^2_{x\pi}.\end{aligned} \tag{A25}$$

However, the excess log return under the objective expectation (which is available to an econometrician) is

$$\begin{aligned}&E_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}Var_t(rx^{\$}_{t\to t+1,n}) \\ &= \tilde{E}_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}\tilde{Var}_t(rx^{\$}_{t\to t+1,n}) + E_t(rx^{\$}_{t\to t+1,n}) - \tilde{E}_t(rx^{\$}_{t\to t+1,n}) \\ &= \tilde{E}_t(rx^{\$}_{t\to t+1,n}) + \frac{1}{2}\tilde{Var}_t(rx^{\$}_{t\to t+1,n}) - B^{\$}_{\pi,n-1}\left[E_t(\tilde{x}_{\pi,t+1}) - \tilde{E}_t(\tilde{x}_{\pi,t+1})\right] \\ &= -\lambda^{\$}_{xc}B^{\$}_{c,n-1}\sigma^2_{xc} - (1-\lambda)\lambda^{\$}_{x\pi}B^{\$}_{\pi,n-1}\sigma^2_{x\pi} - (1-\lambda)\rho_\pi B^{\$}_{\pi,n-1}(x_{\pi,t} - \tilde{x}_{\pi,t}),\end{aligned} \tag{A26}$$

where the expectation operator $E_t(\cdot)$ denotes the objective expectation. This equation shows that the excess log return under the objective expectation equals the excess log return under the subjective expectation plus an additional term capturing belief biases about the state variables. Specifically, sticky inflation expectations induce biases in inflation beliefs. Recent inflation news positively predicts subsequent forecast errors ($x_{\pi,t} > \tilde{x}_{\pi,t}$), and consequently, negatively predicts future bond excess returns as investors gradually correct their belief errors. In contrast, without sticky expectations (i.e., when $\lambda = 0$), $x_{\pi,t} = \tilde{x}_{\pi,t}$ and bond excess returns are unpredictable by inflation news.

B.6 Term Premium

I can decompose the nth period yield into expectations of average future short rates and average excess returns over the life of the bond:

$$y^{\$}_{t,n} = \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t y^{\$}_{t+i-1,1} + \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t rx^{\$}_{t+i-1\to t+i,n-i+1}. \tag{A27}$$

The last term on the right is often described as the bond's nominal term premium, denoted as $tp^{\$}_{t,n}$.

Using Equation (A21), I show that the term premium under subjective expectations is constant in this model. Specifically,

$$\begin{aligned} tp^{\$}_{t,n} &= y^{\$}_{t,n} - \frac{1}{n}\sum_{i=1}^{n} \tilde{E}_t y^{\$}_{t+i-1,1} \\ &= \frac{1}{n}\left(B^{\$}_{0,n} + B^{\$}_{c,n} x_{c,t} + B^{\$}_{\pi,n}\tilde{x}_{\pi,t}\right) \\ &\quad - \frac{1}{n}\left(nB^{\$}_{0,1} + B^{\$}_{c,1}\sum_{i=1}^{n}\tilde{E}_t[x_{c,t+i-1}] + B^{\$}_{\pi,1}\sum_{i=1}^{n}\rho_{\pi}^{i-1}\tilde{x}_{\pi,t}\right) \\ &= \frac{1}{n}B^{\$}_{0,n} - B^{\$}_{0,1}. \end{aligned} \tag{A28}$$

This result indicates that under subjective expectations with sticky beliefs, the term premium is constant and unpredictable by inflation news.

B.7 Inflation Variance Ratios: Model-Implied

Equation (A27) decomposes an nth period yield into expectations of average expected future short rates and a term premium component. If I define the ex-ante real rate as the yield of a one-period nominal bond minus expected inflation,

$$r_{f,t} = y^{\$}_{t,1} - \tilde{E}_t \pi_{t+1}. \tag{A29}$$

Furthermore, I can decompose the nominal bond yield as

$$y^{\$}_{t,n} = \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t r_{f,t+i-1} + \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t\pi_{t+i} + tp^{\$}_{t,n}. \tag{A30}$$

From this last equation, innovations in the n maturity yield from $t-1$ to t are equal to the sum of news about ex ante real rates, expected average inflation, and term premium shocks:

$$\epsilon^{(n)}_{y^{\$},t} = \epsilon^{(n)}_{r,t} + \epsilon^{(n)}_{\pi,t} + \epsilon^{(n)}_{tp^{\$},t}, \tag{A31}$$

with

$$\begin{aligned} \epsilon^{(n)}_{y^{\$},t} &= \tilde{E}_t y^{\$}_{t,n} - \tilde{E}_{t-1} y^{\$}_{t,n}, \quad \epsilon^{(n)}_{r,t} = \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t r_{f,t+i-1} - \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_{t-1} r_{f,t+i-1}. \\ \epsilon^{(n)}_{tp^{\$},t} &= \tilde{E}_t tp^{\$}_{t,n} - \tilde{E}_{t-1} tp^{\$}_{t,n}, \quad \epsilon^{(n)}_{\pi,t} = \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t\pi_{t+i} - \frac{1}{n}\sum_{i=1}^{n}\tilde{E}_{t-1}\pi_{t+i}. \end{aligned} \tag{A32}$$

Given the equilibrium solution of bond yields and the assumed process for the inflation rate, I am able to derive analytical expressions for each of these shocks. The average expected ex-ante real rate is:

$$\frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t r_{f,t+i-1} = (B^{\$}_{0,1} - \mu_\pi) + \frac{1}{n}B^{\$}_{c,n}x_{c,t} + \frac{1}{n}\left(B^{\$}_{\pi,n} - \frac{1-\rho^n_\pi}{1-\rho_\pi}\right)\tilde{x}_{\pi,t}. \tag{A33}$$

Thus, the innovations to this component are

$$\epsilon^{(n)}_{r,t} = \frac{1}{n}B^{\$}_{c,n}\sigma_{xc}\eta_{xc,t} + \frac{1}{n}\left(B^{\$}_{\pi,n} - \frac{1-\rho^n_\pi}{1-\rho_\pi}\right)(1-\lambda)(x_{\pi,t} - \rho_\pi\tilde{x}_{\pi,t-1}). \tag{A34}$$

Similarly, expected average inflation is equal to

$$\frac{1}{n}\sum_{i=1}^{n}\tilde{E}_t\pi_{t+i} = \mu_\pi + \frac{1}{n}\frac{1-\rho^n_\pi}{1-\rho_\pi}\tilde{x}_{\pi,t}, \tag{A35}$$

with shocks given by

$$\epsilon^{(n)}_{\pi,t} = \frac{1}{n}\frac{1-\rho^n_\pi}{1-\rho_\pi}(1-\lambda)(x_{\pi,t} - \rho_\pi\tilde{x}_{\pi,t-1}). \tag{A36}$$

Therefore, the variance of inflation news is

$$\mathrm{Var}(\epsilon^{(n)}_{\pi,t}) = \frac{1}{n^2}\left(\frac{1-\rho^n_\pi}{1-\rho_\pi}\right)^2\frac{(1-\lambda)^2\sigma^2_{x\pi}}{1-\lambda^2\rho^2_\pi}. \tag{A37}$$

Furthermore, Equation (A28) shows that shocks to the term premium are zero

$$\epsilon^{(n)}_{tp^{\$},t} = \tilde{E}_t tp^{\$}_{t,n} - \tilde{E}_{t-1}tp^{\$}_{t,n} = 0. \tag{A38}$$

It is straightforward to show that the sum of $\epsilon^{(n)}_{r,t}, \epsilon^{(n)}_{\pi,t}$, and $\epsilon^{(n)}_{tp^{\$},t}$ (given by Equations (A34), (A36), and (A38)) is equal to innovations in the n-period yield:

$$\epsilon^{(n)}_{y^{\$},t} = \frac{1}{n}B^{\$}_{c,n}\sigma_{xc}\eta_{xc,t} + \frac{1}{n}(1-\lambda)B^{\$}_{\pi,n}(x_{\pi,t} - \rho_\pi\tilde{x}_{\pi,t-1}). \tag{A39}$$

This implies that the variance of yield innovation is

$$\mathrm{Var}(\epsilon^{(n)}_{y^{\$},t}) = \frac{1}{n^2}\left[B^{2,\$}_{c,n}\sigma^2_{xc} + \frac{(1-\lambda)^2B^{2,\$}_{\pi,n}\sigma^2_{x\pi}}{1-\lambda^2\rho^2_\pi}\right]. \tag{A40}$$

Finally, the model-implied inflation variance ratios for an econometrician are equal to the

variance of Equation (A36) divided by the variance of Equation (A39):

$$\begin{aligned} \text{Inflation Variance Ratio} &= \frac{\text{Var}\left(\epsilon_{\pi,t}^{(n)}\right)}{\text{Var}\left(\epsilon_{y^{\$},t}^{(n)}\right)} \\ &= \frac{\left(\frac{1-\rho_{\pi}^{n}}{1-\rho_{\pi}}\right)^{2}(1-\lambda)^{2}\sigma_{x\pi}^{2}}{(1-\lambda^{2}\rho_{\pi}^{2})B_{c,n}^{2,\$}\sigma_{xc}^{2}+B_{\pi,n}^{2,\$}(1-\lambda)^{2}\sigma_{x\pi}^{2}}. \end{aligned} \tag{A41}$$

As shown in Equation (A41), when investors have stickier beliefs (larger λ, implying slower updating), the inflation variance ratio is attenuated because inflation news is not immediately impounded into bond yields. In the limit, as λ approaches 1, the numerator and inflation-related component of the denominator vanish and yield innovations are driven entirely by real rate news. Conversely, when $\lambda = 0$ (rational expectations), the full inflation signal is immediately incorporated into yields.