Good Inflation, Bad Inflation, and the Dynamics of Credit Risk

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Abstract

We study how changes in expected inflation affect firm-level credit spreads, and uncover evidence of a time-varying inflation sensitivity. In times of "good inflation," when inflation news is perceived by investors to be positively correlated with real economic growth, movements in expected inflation substantially reduce corporate credit spreads. Meanwhile in times of "bad inflation," these effects are attenuated and the opposite can take place. These dynamics are driven by movements in credit risk premia and naturally arise in an equilibrium asset pricing model with a time-varying inflation-growth covariance and persistent macroeconomic expectations.

Keywords: Inflation Sensitivity, Time Variation, Credit Spreads, Stock-Bond Correlation
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1 Introduction

The post-pandemic surge in global price levels has reignited academic and practitioner interest in the corporate credit market response to inflation. Classic theories of debt deflation (e.g., Fisher (1933)) suggest that higher inflation reduces the real value of interest payments, easing firms' debt burdens, lowering default probabilities, and reducing credit spreads. These theories typically rest on the assumption that real cash flows are, on average, uncorrelated with inflation fluctuations. In reality, recent evidence shows a robust, time-varying correlation between inflation and real economic growth in the U.S. (e.g., David and Veronesi (2013)). From the stagflationary era of the 1970s and early 1980s to the more demand-driven, procyclical inflation regime of the past two decades, investors' perceptions of the inflation-growth link have shifted. This dynamic relationship has meaningful implications not only for credit pricing, but also for real economic activity through firms' capital structure and investment decisions.

This paper provides strong evidence of a time-varying response of corporate credit spreads to movements in inflation expectations. Using firm-level credit default swap (CDS) data, we show that corporate spreads decline more sharply following upward revisions in inflation expectations when inflation is perceived to better signal positive economic growth (i.e., more of a "good inflation" environment). To capture a high frequency relationship between inflation and real growth, we use the stock-bond return correlation, as risk-free bond returns decline with positive inflation news and equity returns rise with real growth (e.g., Campbell, Pflueger, and Viceira (2020)). We find a robust time variation in the credit spread response, which operates primarily through credit risk premia, and is more pronounced for riskier and financially constrained firms.[1] Building on Bansal and Yaron (2004) and Augustin (2018), we rationalize the documented time-varying behavior of credit markets through an equilibrium asset pricing model extended to price defaultable credit securities. The model delivers an inflation sensitivity in corporate credit spreads that is tied to the bond-stock return correlation and highlights the role of persistent growth expectations as an amplification mechanism.

Our empirical strategy is centered around changes in expected inflation triggered by macroeconomic announcements. We measure expectation changes using daily and intraday

[1] There is limited work on these topics in credit risk. To the best of our knowledge, we are among the first to empirically examine how the response of corporate credit risk to inflation expectation movements varies with the macroeconomic environment. Kang and Pflueger (2015) discuss the level effects of inflation volatility and cyclicality on corporate credit spreads, while Bhamra, Dorion, Jeanneret, and Weber (2022) show that higher expected inflation reduces default risk unconditionally. More recently, Bonelli (2025) and Lu, Nozawa, and Song (2025) examine the average inflation sensitivities of corporate bond spreads and returns, respectively, using inflation swaps. Relative to these studies, our paper focuses on the time-varying sensitivity of credit spreads to expected inflation news, identified through macroeconomic announcement days.

movements in five-year inflation swap rates. These market-based contracts reflect longer-term inflation expectations and provide a closer link to the long-duration cash flows present in credit securities.[2] Macroeconomic announcements serve as ideal events for this setting, as (i) they prompt investors to incorporate new information and (ii) inflation swaps display a greater degree of variation on these days.[3]

Using daily changes in inflation swaps on macroeconomic announcement days, we first document that, on average, positive revisions in expected inflation reduce credit spreads. Over the full 2004–2023 sample, a one standard deviation (1σ) increase in expected inflation is associated with a 1 basis point (b.p.) decline in five-year CDS spreads, corresponding to a 12% decline in standard deviation terms.[4] These findings are consistent with the predominance of a "good inflation" regime in the U.S. during this period, as reflected in the negative average stock-bond return correlation observed since 2004.

That said, the market-perceived relationship between inflation news and real growth can experience sudden changes due to rapidly evolving economic conditions. Macroeconomic quantities, which are observable at a lower frequency, cannot fully capture changes in the inflation-growth relationship in real-time. For this reason, we use a well-established and economically motivated proxy – the bond-stock return correlation – which we can conveniently track on a daily basis (e.g., Campbell et al. (2020)). In our sample, the daily three-month bond-stock return correlation averages -0.30, consistent with a "good inflation" environment. However, we also observe episodes where the correlation turns positive, reaching levels comparable to those seen during the stagflationary period of the 1970s. For instance, the correlation rose to approximately 0.50 in late May 2021 and again in December 2022, when, for brief periods of time, market participants were potentially interpreting rising inflation expectations as a signal of weaker future real growth.

Our analysis shows that high-frequency changes in the bond-stock return correlation influence how credit spreads respond to movements in inflation expectations. In our baseline panel regression, we include an interaction term between changes in inflation swap rates on

[2] We use five-year inflation swap rates as a reliable reflection of market inflation expectations. They capture variation more accurately than surveys or statistically-based autoregressive forecasts and align closely with realized inflation Diercks, Campbell, Sharpe, and Soques (2023). At medium to long horizons, swaps carry minimal risk premia and adjust rapidly to new information Bahaj, Czech, Ding, and Reis (2023). Unlike inflation protected Treasury securities (TIPS), whose embedded deflation put option biases breakeven rates, zero-coupon swaps offer an unbiased market-implied view of expected inflation.

[3] We focus on macroeconomic announcements related to the Consumer Price Index (CPI), Producer Price Index (PPI), real Gross Domestic Product (GDP), and Nonfarm Payrolls. We choose this set of announcements as their survey-based surprises are significantly priced in intraday inflation swap markets. Additionally, we confirm in the Appendix that the observed time variation extends beyond announcement days to the full sample period.

[4] This effect becomes more pronounced with wider event windows and holds more strongly for longer-horizon (10-year swaps).

macroeconomic announcement days and the lagged stock-bond return correlation. We find that when the correlation is lower – indicating more of a "good inflation" environment – credit spreads fall more in response to positive inflation news. For instance, when the bond-stock correlation is two standard deviations below its average, a one standard deviation increase in expected inflation leads to a decline in credit spreads that is twice as large as the average effect. An analogous interpretation holds in the other direction.[5]

We also provide novel results that better highlight the manner in which credit spreads respond to inflation news. Using a decomposition similar to the one in Berndt, Douglas, Duffie, and Ferguson (2018), we show that the majority of the credit spread effect operates through the risk premium channel. The expected losses (risk neutral) component, which highly correlates with pure default risk, also displays significant time-variation in its inflation sensitivity, but to a lesser degree. These results suggest that investor sentiment interacts with movements in default risk to generate a sizeable reaction in credit spreads through risk premia. Additionally, we show in the cross section that our findings are strongest for riskier firms, as there is a strong interaction between time variation and heterogeneity in inflation responsiveness. When we sort firms according to their CDS spreads the day before the macroeconomic announcement, we find that firms in the highest credit risk group display a sensitivity to movements in inflation expectations that is more than ten times larger than the sensitivity of firms in the lowest risk group.

To obtain a deeper understanding of the inflation sensitivity in credit markets, we decompose our key inflation swap measure in multiple ways. First, we study whether inflation sensitivities in credit spreads are driven by macroeconomic news or a non-surprise component in expectations. The motivation behind this exercise is natural. Macroeconomic surprises fundamentally affect inflation expectations (e.g., Bauer (2015), Binder (2021)).[6] However, movements in inflation expectations are only partly explained by these surprises, suggesting a strong residual component. This residual captures the multidimensionality of macroeconomic announcements—such as context, tone, and accompanying details—not reflected in the headline surprise (e.g., Gürkaynak, Kısacıkoğlu, and Wright (2020)). Using multiple macroeconomic surprises to disentangle daily inflation swap movements into surprise and residual components, we find that the latter is more influential in contributing to the time-varying inflation sensitivity.

[5] While not the focus of the paper, we can also show that qualitatively similar findings hold in equity markets. As shown in the Appendix, equity prices increase at a greater rate following a revision in inflation expectations, when the stock-bond return correlation is lower. These results are somewhat related to those in Boons, Duarte, de Roon, and Szymanowska (2020), where the authors provide evidence of a time-varying inflation risk premium that is linked to the nominal-real covariance.

[6] Surprises denote the difference between realized macroeconomic measures and the median economist survey taken shortly prior to the announcement day.

One caveat to this analysis is that daily inflation swap movements do not precisely measure the change in expectations surrounding the macro surprise. To correct for this issue, we also collect and study high frequency, 60-minute changes in inflation swaps around announcements. To understand which of the two sources (surprise vs. residual) matter for the credit market reaction, we decompose the intraday movements of inflation swaps using a heteroskedasticity-based approach (e.g., Rigobon and Sack (2004), Gürkaynak et al. (2020)), and identify a latent factor that is orthogonal to headline surprises. This latent factor captures the non-headline component of macroeconomic announcements and accounts for over 60 percent of the total variation in the intraday five-year swap change. We show that this latent component significantly affects credit spreads above and beyond the headline component and helps account for their time-varying sensitivity to inflation expectations.

An additional channel we study explores to what extent the credit market response is driven by revisions in inflation expectations. As inflation swaps technically reflect risk neutral inflation expectations, inflation risk premia might influence the time-varying sensitivities. To address this concern, we test multiple proxies for physical inflation expectations that are embedded in inflation swaps. First, we use the model-based estimates from D'Amico, Kim, and Wei (2018) derived from TIPS yields and survey forecasts of inflation. Second, we apply a principal component analysis that captures expected inflation partly through its well-documented negative relationship with real yields (e.g., Ang, Bekaert, and Wei (2008)). In both cases, we find that our main results are fundamentally expectations driven: the time-varying sensitivity of credit spreads is primarily driven by changes in the expected inflation component.

A final issue we tackle in our empirical analysis is the use of the bond-stock correlation as a conditioning variable. While intuitive and timely, this measure may imperfectly proxy for the inflation-growth relationship, as it can also reflect other market dynamics – such as flight-to-quality behavior – embedded in noisy asset prices. To alleviate this concern, we first compare the bond-stock correlation with more direct, lower-frequency measures of the nominal-real covariance based on macroeconomic data. Although these macroeconomic measures provide useful intuition about inflation cyclicality, the high-frequency bond-stock correlation consistently outperforms in many horse race tests. In addition, recent work has highlighted the role of convenience yields in U.S. Treasuries (e.g., Krishnamurthy and Vissing-Jorgensen (2012); Fleckenstein, Longstaff, and Lustig (2014)). Through a simple decomposition, we split our bond-stock correlation into a component associated with convenience yield and another that reflects a "frictionless" risk-free rate, which aligns more closely with our object of interest. While the convenience yield component plays a role during crises, we show that the frictionless portion is the main driver of the time-varying sensitivity of asset prices to inflation expectations.[7]

In the last part of the paper, we rationalize our empirical analysis using an economic model that features a time-varying nominal-real covariance and persistent macroeconomic expectations (i.e., long-run risks). Building on Bansal and Yaron (2004) and Bansal and Shaliastovich (2012), this time-varying covariance between expected real growth and inflation shocks determines the good and bad nature of expected inflation movements. While shocks to expected inflation raise discount rates in all regimes, they affect a firm's real cash flows and asset prices in an asymmetric manner. To map the model to credit spreads, we also extend the long-run risks framework to price defaultable CDS, while accounting for the time-varying dynamics. This extension yields a numerically tractable, nonlinear expression for CDS spreads that is a function of model parameters, growth and inflation expectations, and the real-nominal covariance regime.

The calibrated model provides a number of implications consistent with the empirical analysis. First, the endogenous model-implied bond-stock return correlation behaves one-to-one with the real growth-inflation covariance. While this covariance ideally would be available on a real-time basis in the data, our framework shows that the bond-stock correlation serves as an excellent proxy. Second, when the covariance is significantly positive (i.e., a good inflation regime), the model displays credit spreads that negatively respond to expected inflation shocks. Finally, our model speaks to the importance of persistent expectations. When the long-run mechanism in expected growth is attenuated, the bond-stock correlation becomes less volatile, and expected inflation shocks are less relevant for asset prices on an absolute basis.[8]

Related Literature. Our paper relates to a broad set of economic research studying the reaction of financial markets to macroeconomic news, the state dependency in this response, and structural models designed to examine how inflation news in particular affects equity and credit markets.

While a large strand of the high-frequency asset pricing literature has focused on the transmission of monetary policy shocks measured over a narrow window (e.g., Bernanke and Kuttner (2005), Gürkaynak, Sack, and Swanson (2004)), more recent papers have focused on inflation surprises. Gil de Rubio Cruz, Osambela, Palazzo, Palomino, and Suarez (2022)

[7] In the Appendix, we show that our results remain robust to multiple tests, including (among others) the removal of low CDS liquidity observations, using breakeven inflation in place of inflation swaps, and replacing the bond–stock correlation with a measure based on inflation swaps and market returns.

[8] The link that our model draws between the real-nominal covariance and the bond-stock correlation is similar to the New-Keynesian model discussion in Cieslak and Pflueger (2023). Meanwhile, Chernov, Lochstoer, and Song (2021) and Jones and Pyun (2024) study the role of consumption growth persistence toward the volatility of the bond-stock correlation.

show that firm-level close-to-open equity returns react negatively to core CPI surprises, and that firm-level characteristics (e.g., market beta, leverage, and firm size) matter for the transmission. Knox and Timmer (2023) also show that stock prices decline following a positive inflation surprise, more so for firms with low market power. Chaudhary and Marrow (2022) focus on one-day movements in inflation swaps surrounding CPI announcements and show that increases in swap-implied inflation expectations increase equity prices. Closer to our work, some papers have examined the unconditional link between inflation and corporate credit risk. Bhamra et al. (2022) show that higher expected inflation reduces default risk and equity prices simultaneously. More recently, Bonelli (2025) and Lu et al. (2025) examine the average inflation sensitivities of corporate bond spreads and returns, respectively, using inflation swaps. Relative to these papers, we focus on the time-variation of inflation sensitivities, present in credit markets via credit default swaps.

Recent papers have also studied the state dependent pricing of macroeconomic risks, predominantly in equity markets. Elenev, Law, Song, and Yaron (2024) use an array of macroeconomic announcements (capacity utilization, nonfarm payrolls, CPI, GDP, among others) to show that stock markets react more steeply when the output gap is higher and short-term rates are expected to increase. An early paper that discusses the state-dependency in the pricing of CPI surprises is Knif, Kolari, and Pynnönen (2008), where the authors characterize the response of monthly equity prices to CPI surprises, as a function of underlying manufacturing capacity utilization. Similarly, Gil de Rubio Cruz et al. (2022) show that the stock market sensitivity to inflation surprises is the largest during periods when short-run inflation expectations are above their long-run values and the real economy is overheating. Boons et al. (2020) show that the time-varying covariance between inflation and future consumption growth helps determine the equity-implied inflation risk premium. Another recent paper that studies state dependency with respect to inflation news is Kroner (2023), who shows that the transmission of inflation surprises into risk-free bond yields is higher when investors pay greater attention to inflation news. Our study is fundamentally different from these one as we focus on credit market behavior, risk premia in these markets, and on a higher frequency proxy of the nominal-real covariance, the stock-bond correlation.[9]

Structural models of asset prices have also examined the effect of a time-varying nominal-real covariance. Building on the long-run-risks frameworks of Bansal and Shaliastovich

[9]The use of this correlation measure as a financial market indicator of the relationship between real growth and expected inflation goes back at least to Hasseltoft and Burkhardt (2012) and is subsequently studied in David and Veronesi (2013) and Campbell et al. (2020). While some papers have suggested that bond yields and their correlation with stock returns are not significantly indicative of expected inflation news (e.g., Duffee (2018), Duffee (2022)) we provide additional evidence that the correlation measure behaves similarly to fundamentals-based measures, when examining the time-variation in inflation sensitivities.

(2012), Hasseltoft and Burkhardt (2012), and Song (2017), we embed a regime-switching relationship between shocks to expected growth and expected inflation and provide evidence consistent with a real-nominal covariance shift in the early 2000's. Our setting departs from these models by pricing credit default swaps, which lets us trace time-varying inflation sensitivities in credit spreads. Similar to the regime-switching covariance in our model, Kang and Pflueger (2015) highlight the importance of the cyclicality of inflation shocks towards the level of credit spreads in the context of a real business cycle model. Whereas Boons et al. (2020) generate a time-varying equity inflation-risk premium by allowing past inflation shocks to influence future consumption growth, we show that adding persistent long-run expectations amplifies the nominal–real covariance and carries this amplification into credit risk pricing. Finally, unlike the sticky-leverage model of Bhamra et al. (2022), which predicts narrower spreads when expected inflation rises, and the model of Gomes, Jermann, and Schmid (2016), which links disinflation under nominal debt to wider spreads and weaker economic activity, our model features episodes in which spreads widen following increases in inflation expectations and explains them with a mechanism in which real cash flows can become negatively exposed to inflation shocks.

In what follows, Section 2 provides details regarding the key data used in our study, while Section 3 focuses on our empirical tests. In Section 4, we discuss a model that rationalizes our empirical analysis.

2 Data

This section describes the main data used to investigate the response of financial markets to changes in inflation expectations. The key objects of interest are inflation swap spreads, firm-level corporate CDS spreads, and the time-varying correlation between aggregate stock and Treasury bond returns. All are available daily and our sample period is from August 2004 to October 2023, with a focus on behavior on major macroeconomic announcement days. We also use intraday inflation swap prices to confirm that daily patterns hold in a more precise setting.

Inflation Swaps. Inflation swaps are traded instruments that reflect expectations of future inflation. Each swap involves two cash flow components: a fixed payment set at the contract rate and a floating payment tied to the realized CPI inflation over the contract's term. By no-arbitrage, the contract rate represents "expected inflation;" however, as a traded security with future payoffs, it also includes a risk premium.[10]

Inflation swaps are useful for our study in a number of ways. As market contracts for longer-term inflation expectations, they link asset prices with longer-duration cash flows to the relevant views of market participants. This is different than looking at CPI inflation surprises, which are backward-looking, or inflation surveys like the Survey of Professional Forecasters or the Blue Chip Economic Indicators survey, which are only updated monthly or quarterly and do not capture immediate investor views. To this point, Diercks et al. (2023) show that inflation swaps provide better forecasts of future inflation than survey-based measures. Moreover, breakeven inflation implied by TIPS offers a comparable measure of inflation expectations, but using swaps helps to avoid some of the liquidity issues prevalent in TIPS markets (see, e.g., Fleming and Sporn (2013), D'Amico et al. (2018)).

We use daily swap spreads from Bloomberg, focusing on the five-year horizon to align with the maturity of our CDS data. We also study higher frequency inflation swap prices to capture precise movements in expected inflation surrounding macroeconomic release times. These data are collected through Refinitiv Tick History and are available on a minute-by-minute basis going back to October 2007. As all of the key announcements occur at 8:30 AM ET, we compute intraday swap price changes in a 60-minute window (15 minute before and 45 minutes after), similar to the wide window shock in Gürkaynak et al. (2004). Across all 622 macroeconomic releases, changes in five-year inflation swaps display a volatility of roughly 3.3 bps over the announcement window. This is fairly large considering that the daily counterpart displays a standard deviation of 4.9 bps.

Corporate CDS. To proxy for firm-level corporate credit risk, we use single-name CDS spreads at the five-year maturity from Markit. Our approach closely follows Berndt et al. (2018), focusing on senior, unsecured bonds (tier category SNRFOR) with a no-restructuring (XR) clause, and excluding data from the *Financials, Utilities*, and *Government* sectors. Overall, our sample contains roughly 1400 unique firms which shrinks once we focus on firms with EDF data. To control for outlier values, we winsorize all data at the 0.5 percent level. The average five-year CDS spread in our sample is 2.26 percent and exhibits a significant degree of skewness and kurtosis. The daily change in CDS spreads on macroeconomic announcement days displays notable variation (8.4 basis points).

Relative to corporate bonds, there are multiple reasons why CDS data are ideal for our

[10]We recognize that this latter inflation risk premium might be non-trivial and time-varying, however Bahaj et al. (2023) use transaction-level data of traded UK inflation swaps to show that the supply of long-horizon inflation protection is very elastic, reflects fundamentals, and incorporates new information quickly. That said, in Section 3.2.2, using estimates from D'Amico et al. (2018) and a separate PCA analysis, we show that the large majority of the time-varying sensitivity with respect to inflation compensation is driven by physical inflation expectations.

study. First, since CDS are insurance contracts tied to default events of firms, they reflect a risk spread that does not depend on the choice of a risk-free rate. Second, because CDS contracts are traded frequently by a number of institutions (hedge funds, banks, insurance companies, etc.) relative to corporate bonds that trade infrequently, they are less susceptible to pricing frictions that arise from illiquidity and imperfect information (see Bai and Collin-Dufresne (2018)). Finally, a longer-standing literature suggests that CDS lead corporate bonds in price efficiency, which is relevant when we think of the responsiveness of asset prices to inflation news (e.g., Blanco, Brennan, and Marsh (2005), Lee, Naranjo, and Velioglu (2018)).

Stock-Bond Correlation. Our analysis focuses on time-variation in the inflation sensitivity of credit and equity markets and its connection to fundamental economic factors, particularly the *inflation-growth* relationship. A precise measure of this object would help us understand whether inflation movements are the result of positive real growth ("good inflation") or might harm real activity in the future ("bad inflation"). As Cieslak and Pflueger (2023) suggest in different language, inflation can be supply-driven, as it was in the second half of the 20th century, or demand-driven, as it has been more recently.

To approximate the inflation-growth relationship, we use the correlation between stock and U.S. Treasury bond returns. While not a "pure" indicator of inflation and growth, the bond-stock correlation serves as a good proxy. In Figure 1, we show rolling three-month (3M) and six-month (6M) correlations of daily aggregate stock returns (from Ken French's database) and daily U.S. Treasury bond returns (using zero-coupon 5-year yields). As is well documented in other studies, the stock-bond return correlation was strongly positive until the late 1990s, then shifted to a predominantly negative regime. This trend is clearly shown in the second panel of Figure 1, where we focus on data from July 2004 onward, overlapping with our inflation swap sample. Although there have been brief periods of positive correlation over the past 20 years (e.g., the mid-2000s and the past two years), the overall trend points to a shift from bad to good inflation regimes.

Despite the shift toward a good inflation regime, our data still show considerable variation in correlation measures. For instance, the average three-month correlation in the shorter sample ranges from -78 percent to 54 percent. In what follows, we exploit this variation to explore how the inflation-growth relationship amplifies effects on asset prices.

3 Empirical Results

Our empirical design relies on the information revealed by macroeconomic announcements. Specifically, we focus on days with data releases related to key price movements (CPI and PPI) or economic activity (nonfarm payroll and initial real GDP release). Market participants respond to surprises in these announcements by adjusting their inflation expectations. By a simple measure, the variance in swap movements is 2 to 3.5 times higher on announcement days compared to non-announcement days.[11]

We start by examining daily changes in credit risk and relate them to movements in swap rates. Our baseline specification is:

$$\Delta s_{it} = \beta_i + \beta_\pi \Delta\pi_t^{swap} + \beta_X' X_{i,t-1} + \varepsilon_{it}, \tag{1}$$

where Δs_{it} reflects the one-day change in CDS spreads ($\Delta s_{it} \equiv s_{it} - s_{i,t-1}$). $\Delta\pi^{swap}$ is the one-day change in five-year swap rates. Lagged variables ($X_{i,t-1}$) include CDS spreads as these might also mechanically affect the daily change in spreads and returns. We control for firm fixed effects (β_i) and cluster standard error by firm-date, as there might be greater comovement of asset prices on event days.

The results for this regression are reported in column (1) of Table 1. We find that a positive change in inflation swaps significantly reduces CDS spreads – that is, higher expected inflation unconditionally reduces credit risk. A one standard deviation change in inflation swaps is associated with a 0.90 basis point reduction in CDS, which corresponds to about 12 percent of the daily standard deviation in CDS rate changes during relevant macroeconomic announcement days.[12] Positive movements in inflation swaps are good news for firms, as CDS spreads decline. These results are also consistent with the average negative stock-bond return correlation in our sample, which indicates a good inflation regime.

3.1 Inflation Cyclicality and Credit Risk Transmission

This unconditional effect may mask time variation and potential reversals. Our hypothesis is that the market-perceived relationship between expected inflation and growth matters significantly for valuation purposes. To empirically test for this time variation, we add an interaction term between expected inflation changes and the stock-bond return correlation,

[11] The degree of variance differences, between announcement and non-announcement days is dependent on the inflation swap maturity. Using swap prices in a narrow window around typical news release timings, we show that two- and three-year swaps display the highest degree of variance increases.

[12] In unreported results we show that the effect further increases in a five-day window following the macroeconomic announcement.

which we use as a proxy for the inflation-growth relationship. We measure this correlation over a 3-month horizon and, for robustness, over a 6-month horizon.[13] The baseline specification in Equation (1) becomes:

$$\Delta s_{it} = \beta_i + \beta_\pi \Delta\pi_t^{swap} + \beta_\rho \tilde{\rho}_{t-1} + \beta_{\rho\pi} \left(\tilde{\rho}_{t-1} \times \Delta\pi_t^{swap}\right) + \beta_X' X_{i,t-1} + \varepsilon_{it}, \tag{2}$$

where $\tilde{\rho}$ is one of the correlation measures.[14] We standardize $\tilde{\rho}$ so that $\beta_{\rho\pi}$ indicates the additional sensitivity to changes in inflation swap when $\tilde{\rho}$ is one standard deviation (1σ) higher. Results from this test are displayed in Table 1. In column (2), we show that a standard deviation reduction in $\tilde{\rho}$ (a movement of about 0.28) leads to a 0.61 b.p. larger reduction in credit spreads following an increase in inflation swaps. Column (3) also displays that similar results hold when we use a longer horizon measurement of the stock-bond correlation.

Our results have an intuitive explanation. Because risk-free bonds yield negative real returns during inflationary periods, while stock returns align with longer-term growth expectations, the stock-bond return correlation serves as a negative proxy for the inflation-growth relationship. Lower values of $\tilde{\rho}$ suggest that inflation movements are further seen as "good inflation," while higher, positive values — as observed in the mid-2000s, mid-2010s, and more recently — indicate "bad inflation." Consistently, when $\tilde{\rho}$ is very negative, the CDS response to inflation shocks is more pronounced downward. For example, with $\tilde{\rho} = -2$, the response to $\Delta\pi^{swap}$ becomes $-2.03 = -0.81 - 2 \times 0.61$. Conversely, when $\tilde{\rho}$ is positive, a strong stock-bond correlation can drive up credit risk following an inflation increase.

Overall, our analysis offers direct evidence that credit markets respond in a time-varying manner to revisions in expected inflation. News of rising inflation expectations reduces credit risk more significantly when inflation movements signal a future economic expansion. Conversely, positive inflation expectations reduces credit risk less—or may even increase it—when they signal potential economic slowdowns ahead. We also show in the Appendix that this credit-based response to expected inflation movements is qualitatively consistent with the response in equity returns, suggesting a cohesive story across both asset classes.

3.1.1 Credit Risk Premia

Corporate credit spreads contain information with respect to risk-neutral compensation for default risk ("expected losses") as well as a risk premium component that reflects the co-movement of investor marginal utility and losses in default. While risk premia are more

[13] In the Appendix, we also replace the stock-bond correlation with a measure based on inflation swaps and market returns, which produces similar or stronger results.

[14] It is key that this correlation is taken at the $t-1$ date, so as to ensure that the news ($\Delta\pi_t^{swap}$) is not taken into account in the ex ante measurement.

difficult to measure in equity returns and corporate bond spreads, credit default swaps allow us to more cleanly decompose spreads due to the deterministic nature of cash flows and limited optionality. Using an approximation of the methodology in Berndt et al. (2018), which we describe in detail in the Appendix, we compute expected losses (EL_{it}) and risk premia (RP_{it}) such that $s_{it} = EL_{it}+RP_{it}$. Using both these components, we modify Equation (2) using either ΔEL_{it} or ΔRP_{it} as the dependent variable and additionally controlling for the lagged expected loss component.

Table 2 reports the results. The sample size in these tests shrinks by roughly half, as the measurement of expected losses requires matching Markit to Moody's EDF data. That said, the average sensitivity of five-year CDS changes to expected inflation is roughly equal to the coefficient obtained with the larger sample as reported in column (1). Columns (2) and (3) suggest that the large majority of unconditional inflation sensitivity operates through the risk premium channel.[15] Close to two thirds of the overall sensitivity is attributable to ΔRP. More importantly, columns (5) and (6) show that risk premia drive most of the time-varying effect. Inflation-growth perceptions primarily influence the pricing of inflation expectation movements through risk premia, however, the interaction term is also significant in the expected loss component. As variation in the expected loss component is largely driven by default probabilities, this suggests a more modest yet significant time variation in the sensitivity of default probabilities to expected inflation movements. Columns (7) through (9) confirm the robustness of these findings with a longer stock-bond correlation window.

3.1.2 Time Variation in the Cross-Section

Credit spreads exhibit a great degree of skewness and kurtosis. In particular, firms with low distances to default and greater financial constraints display increased sensitivities to aggregate news. We combine the cross-sectional heterogeneity with time variation to study potential interaction effects. We re-examine the results from Equation (2) by credit risk group, using a simple measure of risk – a quintile sort of CDS spreads on the day *prior* to the macroeconomic announcement. We do not seek to argue that this is the singular measure of firm-level risk, but it does contain some desirable properties vis-à-vis simple accounting-based measures (e.g., Palazzo and Yamarthy (2022)).

We report results in Table 3. In columns (1) through (4) we focus on results with respect to the CDS spread. To facilitate comparisons to the average effect, the first column repeats an earlier result regarding the time-varying nature of inflation responsiveness, focusing on

[15] In theory, coefficients from the ΔEL and ΔRP should add up to those from the overall spread regression. The minor discrepancies in the table arise from winsorization of all firm-level dependent and independent variables.

firm-dates with a risk premium measurement. Meanwhile, the next three columns focus on risk groups 1, 3, and 5 based on the ex-ante CDS values. There are two main takeaways: (a) the average response is amplified in riskier firms and (b) the degree of time-variation increases dramatically for riskier firms. In the right most columns of the same table, we show that risk premia also behave in a similar manner. The time variation in the sensitivity of risk premia to expected inflation movements is most stark for the riskiest of firms (column (8)), relative to safer firms (column (6)).

These findings clearly illustrate the need to jointly think about the cross-section and time-variation of inflation sensitivity. The results in column (2), concerning safe firms, show that CDS spreads decline by $0.52 = 0.20 + 2 \times 0.16$ basis points following an increase in inflation expectations, when the bond-stock correlation is particularly negative ($\tilde{\rho} = -2$). For a relatively risky firm however (group 5), the overall response is more than ten times as large ($-2.21 - 2 \times 1.74 = -5.69$ b.p.).

3.2 Key Drivers of the Inflation Sensitivity

In this subsection, we study the inflation swap measure in greater detail to better understand the sources of inflation sensitivity. As our empirical setup focuses on macroeconomic announcement days, we first examine whether the sensitivity of credit spreads stems from headline surprises, using daily and intraday swap prices. Following this analysis, we then study whether our inflation swap measure truly reflects inflation expectations.

3.2.1 The Role of Macroeconomic Surprises

Macroeconomic surprises fundamentally impact inflation expectations (e.g., Bauer (2015)). For example, a higher than expected CPI release tends to raise forward-looking expectations of inflation. As our empirical environment is particularly focused on such release days, it is natural to ask if credit spreads directly respond to macroeconomic news or if they reflect the endogenous (latent) formation of inflation beliefs following the announcement. To answer these questions, we take two different approaches that yield a similar answer. We first examine evidence from daily inflation swap prices, followed by a related extension using intraday swap data. Overall, we show that the average and time-varying inflation sensitivity in credit spreads is mostly driven by the non-surprise component of inflation expectations.

Non-Surprise Component of Daily Swap Movements. Our first approach focuses on a decomposition of daily movements in inflation swaps. To do so, we use a regression-based exercise, projecting swap changes onto surprises:

$$\Delta\pi_t^{swap} = \gamma_0 + \gamma_s' s_t + \varepsilon_t$$

where s_t indicates a vector of surprises that we study throughout the paper. We only include observations when one of announcement days (CPI / core CPI, PPI / core PPI, nonfarm payrolls, initial real GDP) occurs. On each announcement day, we set all other announcements to zero (i.e., on a CPI / core CPI day, all remaining announcement values are set to 0). For robustness, we also study an alternative price-based set of days, where we only include CPI and PPI days. In unreported results, we find that the R^2 of these regressions is roughly 10–15%, suggesting sizable variation in the residual component of inflation expectation movements to begin with.

Using these results, we can decompose $\Delta\pi^{swap} = \Delta\hat{\pi}^{surp} + \Delta\hat{\pi}^{resid}$, where $\Delta\hat{\pi}^{surp} = \hat{\gamma}_0 + \hat{\gamma}_s' s_t$. In Table 4, we examine which of these two components matter for the time-variation in credit spread sensitivities to expected inflation. Columns (1) and (2) replicate our results from earlier, while columns (3) and (4) display the decomposition. We find that the residual component in inflation expectations plays a crucial role. With respect to the average effect, column (3) shows that the surprise is insignificant while the residual is strongly significant. And finally in column (4), we show that while $\Delta\pi^{surp}$ is indeed priced in a time-varying manner, it is the residual that elicits the strongest time variation (close to two times as much). The results in columns (5) – (8), which focus on price-based announcement days, confirm these results. Overall, the daily data suggest that the non-surprise component of inflation expectation movements is crucial to understand credit spread dynamics.

Time-Variation Using Intraday Swaps. In the previous analysis, the inflation expectation movements were measured using daily changes. One caveat to this exercise is that these movements might not precisely capture the immediate change in expectations upon announcement, due to additional news or factors that might affect inflation swap prices over the course of the day. To remedy these issues, we follow identification techniques used in the high frequency literature (e.g., Gürkaynak et al. (2004)) and focus on inflation swap changes in a 60-minute window (8:15 AM to 9:15 AM EST) around our macroeconomic announcement, all of which are released on a monthly or quarterly basis. To avoid overlap with FOMC announcements, we exclude days when both macroeconomic and FOMC news are released. This results in 622 announcements from the merged sample of intraday inflation swap data, available starting from October 2007.[16]

[16]We provide basic sample statistics and more information regarding the intraday data in the Appendix. For example, we confirm that our chosen macroeconomic announcements are of relevance for movements in

To address our fundamental question regarding the impact of surprise and residual (latent) components of inflation expectations, we take a slightly different approach with the intraday data. Instead of a simple regression as before, we apply Gürkaynak et al. (2020)'s heteroskedasticity-based approach to inflation swaps and use a Kalman filter to identify a latent factor orthogonal to macroeconomic news.

Using intraday inflation swap data from both announcement and non-announcement days, we estimate the following model :

$$\Delta\pi_t^{idswap} = \beta' s_t + \gamma d_t f_t + \eta_t, \tag{3}$$

where $\Delta\pi_t^{idswap}$ is a 6×1 vector of 60-minute window, intraday changes in inflation swaps rates for various maturities (1, 2, 3, 5, 7, and 10 years), and s_t is the vector of macro surprises from before. If an announcement occurs on a given day, d_t equals 1 (otherwise 0) and f_t is an I.I.D. $\mathcal{N}(0,1)$ latent variable that captures the unobserved surprise component. The estimated latent factor is common across maturities, with varying loadings by maturity (i.e., γ is also 6×1). Moreover, γ is identified through the increased variance of swap price movements on macro surprise days, present across multiple swap maturities.[17]

Through our estimation we find that incorporating the latent factor significantly increases explanatory power, allowing us to explain the majority of inflation swap curve movements during announcement dates. Next, we decompose intraday changes in inflation swaps into headline (surprise) and non-headline (latent factor) components,

$$\Delta\pi_t^{idswap,i} = \underbrace{\beta_i' s_t}_{\Delta\pi_t^{surp,i}} + \underbrace{\gamma_i d_t f_t}_{\Delta\pi_t^{latent,i}} + \eta_t^i, \tag{4}$$

where we have rewritten the earlier equation for a specific maturity i. Focusing on the five-year maturity, we modify our baseline regression to include both components:

$$\begin{aligned}\Delta s_{it} = \beta_i + \beta_{\pi_s}\Delta\pi_t^{surp} + \beta_{\pi_l}\Delta\pi_t^{latent} + \beta_\rho\tilde{\rho}_{t-1} + \\ \beta_{\rho\pi_s}\left(\tilde{\rho}_{t-1}\times\Delta\pi_t^{surp}\right) + \beta_{\rho\pi_l}\left(\tilde{\rho}_{t-1}\times\Delta\pi_t^{latent}\right) + \beta_X' X_{i,t-1} + \varepsilon_{it}.\end{aligned} \tag{5}$$

Table 5 presents these regression results. In columns (1) and (2) we confirm that the baseline

inflation swap prices within the 60-minute window.

[17] More precisely, the Gürkaynak et al. methodology requires swap residuals to be heteroskedastic, with larger residual variance on announcement days compared to non-announcement days. We demonstrate this statistically and provide other details regarding the estimation in the Appendix. We thank Gürkaynak et al. for graciously making their Kalman filter code available to the public. While the original application in their paper involves identifying a latent factor in high-frequency interest rate and equity futures, we adapt their code to an inflation swap setting.

daily-based swap results hold in a limited sample where we have intraday data (post-2007). In columns (3) and (4) we replace daily movements in swap rates with high-frequency movements ($\Delta\pi_t^{idswap}$). While the magnitudes are lower relative to the daily data, column (4) in particular shows that CDS spreads continue to display a time-varying sensitivity to intraday spread changes, as a function of the recent bond-stock correlation. We interpret this finding as markets being slower to react; the response becomes clearer over the course of the trading day as inflation swaps further incorporate economic information.

In column (5), we show the response of CDS spread changes to the surprise and latent factor components. Consistent with the estimation results, the latent factor has the largest effect and drives the negative reaction of CDS spreads unconditionallly. In column (6), we incorporate interaction effects with the lagged bond-stock return correlation and find significant time-varying sensitivities, for both the surprise and latent factor components. Similar to the results using daily swap prices, the non-surprise factor contributes more significantly to the time-variation.

3.2.2 Inflation Expectations Versus Inflation Risk Premia

As inflation swap rates (and breakeven inflation yields) measure expected inflation under the risk-neutral measure, they may contain a risk premium component. Hence, interpreting them as a true expected inflation measure can be difficult. In this subsection, we examine whether our results are mostly driven by movements in expected inflation or risk premia. To decompose any inflation compensation measure, an underlying model is required to inform expectations, which could potentially be subject to misspecification. To gain a robust understanding of the role of inflation expectations, we study multiple decompositions – one that is based on the term structure model of D'Amico et al. (2018) and another that is more statistical in nature and based on a principal component analysis of relevant data series.

D'Amico et al. (2018) estimate an affine term-structure model fitted primarily to daily data on nominal yields, TIPS yields, and survey inflation forecasts. Through an estimation procedure, they identify a component of breakeven inflation that results from TIPS illiquidity, and are able to disentangle the residual inflation compensation measure between inflation expectations and inflation risk premia. Using their daily data, we replace our swap measure with these two measures. Results are presented in Table 6. In columns (1) and (2) of the table we separately test their measures of expected inflation and inflation risk premium at the five-year horizon with respect to movements in CDS spreads. While the two measures are certainly correlated, we find that expected inflation drives a greater share of average and time-varying inflation sensitivity.

As a robustness check on these results, we also conduct a principal component analysis of relevant data series in order to decompose inflation swap rates. Intuitively, we know that both inflation swap rates and breakeven yields provide measures of inflation compensation. To help identify the expected inflation component we also include real bond yields, which have been shown to correlate negatively with expected inflation measures (e.g., Pennacchi (1991), Kandel, Ofer, and Sarig (1996), Ang et al. (2008)). Using these three series (inflation swaps, breakeven inflation, and TIPS yields), we estimate a PCA of daily changes going back to 2004. The first two principal components of the PCA help explain a large portion of the variance (greater than 90%). As shown in the Appendix, the first principal component loads negatively on real rates (and positively on the other two series), while the second principal component loads positively on all three data series. Consistent with the prior literature, we label the first principal component as an expected inflation (ΔEI) factor and the second as a risk premium (ΔRP) factor.[18] An additional check outside the PCA, confirms that our cumulated EI factor correlates positively and significantly with 1-year and 5-year ahead inflation, above and beyond inflation expectations from the Michigan Survey of Consumers.[19]

Using these two factors we decompose inflation swap movements into an expected and inflation risk premium component through a simple regression:

$$\Delta\pi_t^{swap5Y} = \gamma_0 + \gamma_E \Delta EI_t + \gamma_R \Delta RP_t + \eta_t.$$

Taking each fitted value separately from this regression ($\hat{\gamma}_E \Delta EI_t, \hat{\gamma}_R \Delta RP_t$), we examine which component matters more for time-varying inflation sensitivites. In columns (3) and (4) of Table 6, we show in CDS markets that the expected inflation component is significantly more important than risk premia. The former elicits a stronger conditional and unconditional response and delivers a higher adjusted R^2 value.

In summary, our results using the D'Amico et al. (2018) and PCA-based measures suggest that expected inflation is a key driver of inflation swaps. Furthermore, the expected inflation component plays a pivotal role in generating time-variation in the inflation sensitivites of asset prices.

3.3 Interpreting the Bond-Stock Correlation

Our baseline results show that movements in expected inflation affect credit and equity markets in a robust, time-varying manner, as dictated by the bond-stock return correlation. However, there are important questions that remain regarding the interpretation of the

[18] In theory it should not be the case that expected inflation and inflation risk premia are unrelated, while they are mechanically independent using the PCA approach. One way to interpret our procedure is that our second factor is an orthogonalized version of the inflation risk premium.

[19] Additional details and analysis can be found in Appendix A.

correlation. For example, as the correlation measure is based purely on financial market prices can it compare well with more fundamentals-based measures that directly measure the nominal-real covariance? Also, recent literature on convenience yields in the U.S. Treasury market emphasizes an outsized role of flight-to-safety effects and market noise. How is the bond-stock correlation and the inflation sensitivity of credit spreads influenced by these effects? We explore both of these issues in this subsection.

3.3.1 Alternative Measures of the Nominal-Real Covariance

Our baseline specification uses the stock-bond correlation as a proxy for the time-varying nominal-real covariance. While our model – discussed in detail later – establishes a direct link between the two, the bond-stock correlation may be a noisy proxy as it is based on financial market prices. Here, we examine whether monthly nominal-real covariance measures, based on macroeconomic data and used in Boons et al. (2020), serves as a better proxy. Additionally, building on Elenev et al. (2024), we test whether capacity utilization, a measure of economic slack, is relevant for the time-varying inflation sensitivity of financial markets.

Similar to the prior literature, we construct a regression-based measure of the real-nominal covariance. We use monthly nondurables and services consumption data from NIPA, deflate it with the PCE price index, and normalize it by the U.S. population time series to create a real, per capita consumption series. We run the following predictive regression:

$$\Delta C_{s+1:s+12} = \alpha_t + \beta_t \Pi_s + e_{s+1:s+12}, \quad \text{for } s = 1, \ldots, t-12, \tag{6}$$

where Π_s is the monthly PCE inflation rate, and $\Delta C_{s+1:s+12}$ is the future annual consumption growth rate. The baseline specification uses an expanding window, weighted least squares, with more weight on recent observations (exponentially decaying with a half-life of 60 months). We also test a rolling regression (OLS) over the past 60 months in our panel regression analysis.

In Table 7, we repeat our baseline tests from Equation 2, replacing the bond-stock correlation with the measures described above. All measures are standardized within the interaction effect term, using values available before the announcement day. Columns (2) and (3) show that the nominal-real covariance measures (expanding and rolling) yield intuitive coefficients: when the covariance is more positive (indicating a good inflation environment), CDS spreads decrease further in response to swap movements. Although the capacity utilization coefficient sign is reasonable, implying that more economic slack leads to a better credit outcome, it is statistically insignificant. In columns (5) through (7), we compare the bond-stock correlation measure with the other three variables through a horse-race of interaction effects. We show that the bond-stock correlation remains the strongest driver of the time-varying inflation responsiveness.

3.3.2 Controlling for Convenience Yields

Due to their (i) cash-like properties, (ii) use as collateral in short-term funding markets, and (iii) store of value for foreign investors, U.S. Treasuries provide additional benefits beyond their pure investment value. Recent literature explore these ideas and show that U.S. Treasuries create additional value (a "convenience yield"), above and beyond other securities that provide identical or similar payoffs.[20] Here, we explore how the convenience yield affects the bond-stock correlation and show that the large majority of our inflation sensitivity results are driven by a "frictionless" risk-free rate that nets out the convenience yield.

As discussed in Acharya and Laarits (2025), we can decompose the nominal Treasury yield at maturity n into three components:

$$yield^n = yield^{*,n} + CDS^{US,n} - convenyield^n,$$

where the yield will contain a frictionless component ($yield^*$), a default risk premium given through the U.S. CDS spread (CDS^{US}) and a convenience yield ($convenyield$).[21] The latter indicates the yield discount, or price premium, that investors are willing to pay to hold Treasuries, beyond the other two components. Rewriting the above equation in terms of returns, based on a log approximation, we receive $R^{bond,n} = R^{bond*,n} + R^{CDS,n} - R^{cy,n}$. Taking covariances with stock returns and dividing through by standard deviations we obtain:

$$\rho^{bond-mkt} = \underbrace{w_1 \rho^{bond^*-mkt}}_{\text{Frictionless}} + \underbrace{w_2 \rho^{CDS-mkt}}_{\text{Default}} - \underbrace{w_3 \rho^{conyld-mkt}}_{\text{Convenience}}, \quad (7)$$

where the weights on the right hand side reflect the ratio of each return's standard deviation to the nominal bond return standard deviation. Put differently, we decompose the bond-stock correlation into three parts related to the frictionless risk-free, a default component, and a convenience yield portion. To better understand how our baseline results arise, we test these components separately, using a well-known measure of convenience yield.

To operationalize this decomposition we fix the maturity of all securities at the 5-year horizon, and correlations to reflect 66-day rolling windows, identical to our baseline specification. To measure the convenience yields, we follow Fleckenstein et al. (2014) and use the

[20]Important and recent papers in this literature include Krishnamurthy and Vissing-Jorgensen (2012), Fleckenstein et al. (2014), Cieslak, Li, and Pflueger (2024) and Acharya and Laarits (2025).

[21]Note that we include CDS^{US} for completeness sake, and do not seek to take a stance on whether there is true default risk in U.S. Treasuries. Our results hold regardless of whether we allow this term to be nonzero in the data.

difference between a synthetic nominal Treasury and the traded Treasury:

$$convenyield^n = \pi^{swap,n} + yield^{real,n} - yield^n,$$

where the sum of the first two terms above (inflation swap rate and the TIPS yield) indicates the yield on a synthetic security which replicates a nominal Treasury's payoffs.[22] The difference between this yield and an actual Treasury should be zero, in a frictionless no-arbitrage setting. However, in practice it tends to be non-zero and fluctuates.

In panel (a) of Figure 2, we plot this measure, also known as the TIPS-Treasury premium. This premium increases dramatically during the Global Financial Crisis (GFC), suggesting a very strong relative premium for Treasuries at this point in time, while afterwards it oscillates consistently around its average value.[23] In the second panel of the same figure, we display multiple measures of the frictionless risk-free rate, in comparison to the standard nominal Treasury. The orange line assumes that there is no default risk (fixing $CDS^{US} = 0$), while the green line incorporates 5Y U.S. CDS spreads from Markit. All three risk-free rates are highly correlated and mostly differ from one another around the GFC. In the final panel, we plot the correlation components from Equation (7), assuming that the CDS component is equal to zero. Overall, the bond-stock correlation is strongly associated with the frictionless component (correlation between the two greater than 90%). Meanwhile, the convenience yield portion has a correlation of -3% as it enters negatively into the decomposition. Generally, $\rho^{conyld-mkt}$ hovers around zero except for a few episodes, surrounding the GFC, shortly after the start of 2018, and in early 2020.

Using these correlation subcomponents, we examine how our overall inflation sensitivity results arise, as shown in Table 8. Column (1) restates our earlier results using the observed bond-stock correlation and changes in CDS spreads. In columns (2) and (3), we test the frictionless bond-stock correlation assuming that U.S. CDS values are equal to zero ($bond^{*ND}$), or are potentially non-zero ($bond^{*D}$) using the U.S. CDS data. In both of these cases, we find that the frictionless bond continues to generate strong results in line with our overall results – declines in the bond-stock correlation lead to more amplified CDS declines. In columns (4) and (5), we horse-race the frictionless measures against the convenience yield results and

[22]The literature on convenience yields is very large and there are many measures to choose from. While we are not rigid on this front, we primarily choose the Fleckenstein et al. (2014) measure as it attempts to connect two securities that provide almost identical cash flows while others (e.g., the AAA-Treasury spread) are not necessarily perfectly alike. Furthermore, Acharya and Laarits (2025) show that this measure works the strongest in relating it to movements in the bond-stock return covariance – essentially the object of our study.

[23]It is worth noting that the average level of the premium is positive both including and excluding the GFC.

we find that indeed the frictionless components are more important.[24]

Full sample results using CDS suggest that while the frictionless piece is important, the convenience yield plays a role as well. However, looking back at the TIPS-Treasury premium in Figure 2, most of the volatility in the premium occurs around the GFC. That is not to say that it does not exist afterwards or before, but it is less volatile and more consistently positive in non-GFC periods. Using this insight, in the final three columns of Table 8 we remove the GFC portion of our sample to test our measures in a more well-behaved sample. In column (7), we find that removing the GFC period and focusing on the frictionless risk-free yield, leads to a sharpening of our baseline results and extremely strong significance for equity markets as well. Finally in column (8), via a horse-race, the convenience yield is shown to be fully insignificant, in comparison to the frictionless yield.

In summary, our results confirm that the convenience yield plays a non-trivial role in nominal Treasury yields. However, as it relates to the time-varying inflation sensitivity of CDS, it is relatively secondary, especially in periods outside the GFC. Furthermore, the fact that our results are driven, and even strengthened, by the frictionless component, confirms our use of the bond-stock correlation as a measure of the inflation-growth relationship.

3.4 Robustness and Extensions

In this subsection, we briefly discuss robustness exercises and extensions. Explicit details regarding each analysis and additional exercises can be found in the Appendix.

Inflation Cyclicality in Equity Markets. The focus of our paper is on corporate credit markets and their sensitivities to expected inflation. However, it is natural to think that equity returns might also display a similar time variation. Indeed, within a matched CDS-equity sample, we show that all our baseline results hold. Equity returns contain a time-varying sensitivity to movements in inflation expectations linked to the stock-bond correlation. Furthermore, magnitudes are quantitatively similar to those in CDS markets. While risk premia measurement is more difficult in equities, relative to CDS, we can relatedly show that the large majority of the time-variation is driven by the systematic component of equity returns. Furthermore, the time variation is exclusively driven by the non-surprise component of inflation expectation movements. These results in equity markets confirm and complement some of the findings in Boons et al. (2020).

Inflation Sensitivities by Swap Horizon. Inflation swaps trade at multiple maturities which might suggest that our results are specific to a particular swap maturity. We show that

[24] In terms of signs, we expect the convenience yield portion to have an opposite sign as it enters negatively into the overall nominal correlation. In this sense, all signs that are reported are intuitive and in line with the baseline results.

our baseline results hold more strongly for longer horizon swaps and attenuate for shorter horizons (the one-year swap). Our results are intuitive from the standpoint that corporate credit spreads reflect the health of long-duration assets subject to longer term fundamentals. They are also consistent with the findings of Bahaj et al. (2023), who show, using regulatory data on inflation swap trading in the U.K., that short term inflation swap prices are mostly informed by liquidity shocks, while fundamentals play a stronger role at longer maturities.

Response to Breakeven Inflation. As mentioned earlier, TIPS-implied breakeven inflation can serve as an alternative measure of expected inflation and strongly relates to inflation swaps. We confirm that our main results hold when using five-year breakeven inflation in place of inflation swaps. Additionally, we find qualitatively consistent results when we replace our bond-stock correlation measure with a measure based on inflation swap movements and market returns.

CDS Liquidity. The Dodd-Frank Act and additional regulations have led to greater standardization and regulation of CDS trading, reducing the gross size of the single-name CDS market over time (e.g., Boyarchenko, Costello, and Shachar (2020)). Consequently, it is important to assess whether our results are affected by low liquidity in CDS markets. To do so, we split CDS observations on each announcement day with respect to the number of underwriting dealers that report quotes. We show that our results actually strengthen for firms with a greater number of participating dealers, suggesting that our results are not driven by low liquidity observations.

Inflation-Growth Regimes over a Long Sample. Although it has turned positive for short periods of time, the bond-stock return correlation has been mostly negative after 2000, making it difficult to detect discrete sign switches in inflation sensitivity. To study discrete sign switches, we extend our credit risk sample back using the aggregate, ICE US Corporate Option-Adjusted Spread index, available from 1996. As an additional exercise, we examine a panel of equity returns back to the 1980s. Using inflation compensation data from D'Amico et al. (2018) surrounding macroeconomic announcements, we show that indeed in periods of negative (positive) correlation, the credit reaction is negative (muted) while the equity reaction is positive (negative). These results suggest that good and bad inflation pricing dynamics are present over a longer time span.

4 Economic Model

In the second part of the paper, we introduce an equilibrium asset pricing model to study how inflation expectations are priced in credit markets. Our model adopts a long-run risks framework (e.g., Bansal and Yaron (2004)), and extends it to provide tractable relationships between credit risk (CDS spreads), the bond-stock correlation, and movements in expected inflation. We present the fundamental aspects and quantitative implications of the model, while reserving solution and calibration details for the Appendix.

A key feature of the model will be a time-varying covariance between shocks to expected inflation and expected real growth. Due to the endowment economy setup, we can specify this relationship exogenously. Obviously, in the real world, there can be deeper fundamentals that determine the covariance (e.g., oil shocks of the 1980's or the COVID-19 pandemic), which we do not seek to characterize here. For our purposes, given a nominal-real covariance, we derive tractable and quantitatively realistic asset pricing implications for the bond-stock correlation and the credit-inflation relationship. The latter might be difficult in more micro-founded models.[25]

Our model explicitly shows that the covariance of expected inflation and real growth is one-to-one with the sign and magnitude of the endogenous bond-stock correlation.[26] Furthermore, the variation in this covariance determines the time-varying sensitivity of risky asset prices to expected inflation news. We conclude by discussing the role of persistent growth expectations, which amplifies the asset pricing effects from shocks to expected inflation.

4.1 Setup

The model is an extension of the long-run risks endowment economy of Bansal and Shaliastovich (2012). We choose a long-run risks framework as the data suggest that movements in expected inflation are crucial determinants of asset prices. Real and nominal fundamentals – that is, consumption growth and inflation – are partially determined by persistent components as follows:

$$
\begin{aligned}
\Delta c_{t+1} &= \mu_c + x_{ct} + \sigma_c \varepsilon_{c,t+1}, \\
\pi_{t+1} &= \mu_\pi + x_{\pi t} + \sigma_\pi \varepsilon_{\pi,t+1}, \\
X_t \equiv \begin{pmatrix} x_{ct} \\ x_{\pi t} \end{pmatrix} &= \Pi X_{t-1} + \Sigma_{t-1}\eta_t, \qquad (8) \\
\Sigma_t &= \begin{pmatrix} \sigma_{xc} & \sigma_{xc\pi}(s_t) \\ 0 & \sigma_{x\pi} \end{pmatrix},
\end{aligned}
$$

[25] As discussed in Campbell et al. (2020) and Cieslak et al. (2024), New-Keynesian models can endogenously generate time-varying nominal-real covariances through shocks to the IS and Phillips curves (producing good and bad inflation, respectively). However, modeling credit securities with default risk and deriving tractable solutions, is more challenging in these environments.

[26] This finding aligns with previous literature explaining the switch in the sign of the bond-stock correlation during the late 1990s. David and Veronesi (2013), Campbell et al. (2020) and Fang, Liu, and Roussanov (2022) attribute this change to the changing correlation between consumption growth and inflation – that is, the nominal channel.

where x_{ct} and $x_{\pi t}$ indicate expected growth and expected inflation, respectively, and the residual components (ε_i) represent short-run noise. Π is the transition matrix for X_t and $\sigma_{xc\pi,t} = \sigma_{xc\pi}(s_t)$ indicates a time-varying parameter that is independently regime-switching and helps determine the covariance of innovations in X_t. The regimes follow an N-state Markov probability matrix, with transition probability from state i to j denoted as p_{ij}. Naturally, $\sum_j p_{ij} = 1$ for all states i.

We intentionally place the regime switching parameter in the composite shock process for growth expectations, as this assumption delivers a direct link between the expected growth level and orthogonalized expected inflation shocks. One can interpret the daily changes in highly persistent inflation swaps as shocks to expected inflation, and this interpretation serves to motivate our setup. However, we are not the first ones to adopt a regime-switching approach as Hasseltoft and Burkhardt (2012) and Song (2017), among others, place regime switches in both the covariance matrix and transition matrix of X_t and estimate these parameters. That said, our goal is to highlight a clear and parsimonious mechanism that works through the expected inflation channel.

In line with the literature, the representative investor has Epstein and Zin (1989) recursive preferences:

$$V_t = \left[(1-\delta)C_t^{\frac{1-\gamma}{\theta}} + \delta\left(\mathbb{E}_t\left(V_{t+1}^{1-\gamma}\right)\right)^{\frac{1}{\theta}}\right]^{\frac{\theta}{1-\gamma}}, \tag{9}$$

where δ is the time discount factor, γ the risk aversion, and ψ the intertemporal elasticity of substitution (IES). The preference for the early resolution of uncertainty is determined by $\theta \equiv \frac{1-\gamma}{1-\frac{1}{\psi}}$. As shown in Epstein and Zin (1989), the investor's (log) pricing kernel takes the form:

$$\begin{aligned} m_{t+1} &= \theta \log\delta - \frac{\theta}{\psi}\Delta c_{t+1} - (1-\theta) r_{c,t+1}, \\ r_{c,t+1} &= \kappa_0 + \kappa_1 pc_{t+1} - pc_t + \Delta c_{t+1}, \end{aligned} \tag{10}$$

where m is the stochastic discount factor, Δc is the log consumption growth, pc is the log price-to-consumption ratio, and r_c is the return on an asset that pays off the aggregate consumption tree as a dividend. Using the Campbell and Shiller (1988) log-linear return approximation, we write the log return in the linear form shown above, where κ_0 and κ_1 are constants that are a function of the average pc. Moreover, for any asset i, including the consumption-paying asset, the Euler condition holds: $E_t\left[\exp\left(m_{t+1} + r_{i,t+1}\right)\right] = 1$.

We focus on the consumption return as a proxy for aggregate equity returns. While the level and volatility of this asset return is less than the empirical counterparts for the aggregate stock market returns, we are mostly concerned about its cyclical properties for the purpose of the bond-stock return correlation. It is straightforward to extend the model to price a levered dividend claim, as done in Bansal and Yaron (2004).

4.1.1 Model Solution and Risk-Free Nominal Bonds

To solve the model, we first characterize the equilibrium price-consumption ratio. Based on the Euler equation and fundamental assumptions, we can show that the price-consumption ratio takes the form $pc_t = A_1' X_t + A_2(s_t)$, where A_1 is a set of loadings on expected growth and inflation and A_2 is a regime-switching component. For a given set of fundamental parameters, A_1 can be solved directly while A_2 is solved numerically through a system of equations.

To compute the bond-stock return correlation, we use both the nominal return on the consumption claim, $r_{c,t+1} + \pi_{t+1}$, and the nominal return on a risk-free bond. The return on an n-period zero-coupon, risk-free bond (purchase at t, sell at $t+1$) is given by:

$$\exp\left(r_{f,t+1}^{\$,n}\right) = \frac{P_{f,t+1}^{\$,n-1}}{P_{f,t}^{\$,n}} = \exp\left(p_{f,t+1}^{\$,n-1} - p_{f,t}^{\$,n}\right), \tag{11}$$

where $P_{f,t}^{\$,n}$ is the price of a nominal risk-free bond at time t maturing at $t+n$, with lowercase indicating its log value. We show that the log price takes the form, $p_{f,t}^{\$,n} = P_1^{n\prime} X_t + P_2^n(s_t)$, where state loadings are maturity specific. Similar to Ang and Piazzesi (2003), we first derive the coefficient values for a one-period risk-free bond and then show that maturity n coefficients can be written recursively. Based on these results, we compute nominal bond prices and corresponding bond returns.

4.1.2 Pricing CDS

We extend the model to price inflation risk in credit markets. While the long-run risks literature largely focuses on asset pricing implications for equity and risk free bond markets, less work has examined its implications for credit markets.[27] As given in Berndt et al. (2018), the CDS of maturity K periods is a rate C_t that satisfies:

$$\Delta C_t \sum_{k=1}^{K/\Delta} \mathbb{E}_t\left[\tilde{M}_{t+k\Delta}^{\$}\left(1 - D_{t,(k-1)\Delta}\right)\right] = \sum_{k=1}^{K/\Delta} \mathbb{E}_t\left[\tilde{M}_{t+k\Delta}^{\$} \times (1-R) \times D_{t+(k-1)\Delta,\Delta}\right], \tag{12}$$

where the left (right) hand side indicates expected payments from the protection holder (seller). Δ denotes the length of time between payments and $\tilde{M}_{t+z}^{\$}$ is the nominal SDF from

[27] Augustin (2018) is an exception and our model uses many elements from his work as a starting point, while embedding the time-varying covariance of real and nominal shocks.

t to $t+z$. $D_{t,z}$ denotes a default indicator between t and $t+z$. For simplicity, we assume constant losses given default $(1-R)$, and that default occurs shortly before the end of each period. Assuming quarterly payments ($\Delta = 1$), we can write the five-year CDS as:

$$C_t = \frac{\sum_{k=1}^{20} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} \times (1-R) \times D_{t+k-1,1} \right]}{\sum_{k=1}^{20} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} (1 - D_{t,k-1}) \right]} = (1-R) \times \left(1 - \frac{\sum_{k=1}^{20} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k} \right]}{\sum_{k=1}^{20} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k-1} \right]} \right), \tag{13}$$

where $S_{t,z}$ indicates a survival dummy variable as of time $t+z$.

Following Augustin (2018) and Doshi, Elkamhi, and Ornthanalai (2018), we assume that default dynamics are exogenous and related to key state variables. While this is a simplification, it allows us to compute CDS prices in closed form and speak to our object of interest – the inflation sensitivity in CDS spreads. Realized default at $t+1$ is given by:

$$D_{t,1} = \begin{cases} 0 & \text{w/probability } \exp(-\lambda_t), \\ 1 & 1 - \exp(-\lambda_t), \end{cases} \tag{14}$$

where the realization is conditionally independent of all other model variables. The ex-ante probability (hazard rate) is based on $\lambda_t = \beta_{\lambda 0}(s_t) + \beta'_{\lambda x} X_t$, which does not guarantee that $\lambda_t > 0$ but it allows us to maintain tractability of CDS prices, given the regime-switching covariance matrix for X. In our quantitative exercise, we ensure a positive λ_t by calibrating $\beta_{\lambda 0}$ and $\beta_{\lambda x}$ appropriately.[28]

To solve for CDS prices, we need to compute, for all k, $\mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k} \right]$ and $\mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k-1} \right]$, under the preference assumptions and model fundamentals. Using the Law of Iterated Expectations and conditional independence assumption of default, we show that there exist coefficients, B_1^k and $B_2^k(i)$, such that $\mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k} \right] = \exp\left(B_1^{k\prime} X_t + B_2^k(s_t) \right)$. Similarly, coefficients C_1^k and $C_2^k(i)$ can be found for $\mathbb{E}_t \left[\tilde{M}^{\$}_{t+k} S_{t,t+k-1} \right]$. These coefficients depend on the fundamental parameters of the model and are solved using a recursive numerical algorithm. Using these results, we can write the model-implied CDS as:

$$C_t = (1-R) \times \left(1 - \frac{\sum_{k=1}^{20} \exp\left(B_1^{k\prime} X_t + B_2^k(s_t) \right)}{\sum_{k=1}^{20} \exp\left(C_1^{k\prime} X_t + C_2^k(s_t) \right)} \right), \tag{15}$$

which is tractable and solves quickly.

[28] One downside of the linear hazard rate formulation is that it restricts the countercylicality of λ_t. To ensure that $\lambda_t > 0$ for $\beta_{\lambda 0}$ and $\beta_{\lambda x \pi} = 0$, we set $\beta_{\lambda xc} > -\frac{\beta_{\lambda 0}}{\max(x_{ct})}$. This limits the volatility of default rates and resulting CDS spreads. Despite this limitation, the model generates reasonable quantitative behavior of CDS spread changes.

4.2 Model Results

In this subsection, we describe the model's key mechanism, illustrate the baseline calibration, and discuss comparative statics. Finally, we show how persistent expectations interact with the time-varying real-nominal covariance.

4.2.1 Key Mechanism

We start by studying the covariance between expected inflation and growth shocks, which is directly affected by $\sigma_{xc\pi}$. We show that this parameter directly connects to the endogenous stock-bond correlation. In this exercise, we assume that $\sigma_{xc\pi}$ is constant and we vary it to examine the model's performance. As we change parameter values, we also ensure that the unconditional variance of expected growth does not change.[29] Fundamental model parameters are set to target values from the data.

Figure 3 displays the model-implied bond-stock correlation based on simulated nominal stock and 5-year bond return data. The y-axis shows the correlation while the x-axis denotes the covariance parameter. Focusing on the solid blue line, the bond-stock return correlation is monotonically decreasing in the covariance. Put differently, when expected inflation shocks are more positively correlated with real consumption growth ($\uparrow \sigma_{xc\pi}$), bond returns become more of a hedge. The reason being that potential shocks to expected inflation increase nominal yields (lead to negative bond returns) while increasing the payoffs of the consumption asset (positive stock returns). A similar interpretation holds in the opposite direction.

Furthermore, this exercise suggests that the model generates sizable movements in the bond-stock correlation. Hence, embeddding movements in $\sigma_{xc\pi}$ can generate plausible variation in this correlation and explain the patterns we see in the data. In what follows, we examine the implications of a time-varying $\sigma_{xc\pi}$.

4.2.2 Model Performance

In line with the long-run risks literature, we calibrate the model at a quarterly frequency and make two simplifying assumptions. First, the autoregressive matrix Π is set to be diagonal with no cross dependencies to allow for a clean interpretation of the covariance parameter as the sole source of the real-nominal interaction. Second, the number of regimes is $N = 2$ so that we have distinctive "good" and "bad" inflation regimes.

Many parameters are taken from the literature or calibrated directly to macroeconomic moments. Regarding the inflation-growth covariance parameter, $\sigma_{xc\pi}$, we calibrate it to be

[29] This is done by directly resizing the constant parameter, σ_{xc} in the growth equation.

positive in the first regime indicating a good inflation regime, and its relative size such that the model unconditionally delivers a negative bond-stock return correlation. Credit-related parameters regarding recovery rates and the sensitivity of default rates to expected growth are also informed by the data.

Based on these parameter values, we solve the model and simulate $40,000$ quarters, including a burn-in period. The first column in Table 9 presents the asset pricing moments of the baseline model. The model does a reasonable job with the annualized nominal risk-free rate (4.63 percent), which is close to the average three-month Treasury bill rate over time. Similarly, the model produces a substantial annual equity premium (0.91 percent) that would be similar to the 5 percent value seen in the data if we employed a levered dividend claim. The average, annualized five-year CDS spread in the model is considerable (1.34 percent) with a reasonable volatility of credit spread movements (5.4 basis points).[30] In terms of risk-free bonds and stocks, the unconditional return correlation in the model is -15 percent as the dynamics from the "good inflation" regime dominate.[31] More specifically, the correlation is -45 percent in the good regime, and 28 percent within the bad regime. These values are reasonable in comparison with the ones documented in the empirical part.

To further understand the model, we use the simulated asset price data to run univariate regressions of spread changes and excess equity returns onto inflation expectation movements. These regressions test the model's ability to generate time-varying asset price sensitivities, similar to the data. In the first column of Table 9, we show that the model can generate differential inflation effects across the two regimes. On average, a standard deviation increase in $\Delta x_{\pi t}$ is associated with a 1.6 basis point decline in CDS spreads. In regime 1, the good regime, this coefficient more than doubles to a 6.3 basis point decline; while in the bad regime, a movement in expected inflation is associated with a 3.1 basis point increase. Similar results obtain for model-implied excess equity returns. A positive movement in inflation expectations increases equity prices. In good inflation regimes, this sensitivity is further amplified. Moreover, across both asset classes, the model displays qualitatively similar behavior as in the data.

[30] The behavior of credit spreads in the model is particularly noteworthy given the parameter restrictions on the hazard rate ($\beta_{\lambda xc}$).

[31] As the correlation is negative, long-term bond returns pay off when the return on consumption is lower and acts as a hedge. As a result, average prices on long-term nominal bonds are higher leading to a negative bond risk premium (i.e. $E\left[r_f^{5Y,\$} - r_f^{\$}\right] < 0$).

4.2.3 Comparative Statics

To better understand the model mechanisms, we examine how the model performs under different parameter configurations and compare them to the baseline.[32] We begin by looking at a calibration where the covariance channel is completely shut off – that is, where $\sigma_{xc\pi} = 0$ across both regimes. This counterfactual helps us determine how much of the asset price response and stock-bond correlation is driven by this channel. Results from this test are presented in the second column of Table 9 ("Model 1"). We see that the absolute size of the bond-stock correlation has shrunk close to zero (0.09) and, similarly, the degree to which risky asset prices respond to inflation shocks is significantly reduced. Now, a standard deviation movement in expected inflation shocks only moves equity returns by about 1 basis point in absolute terms, compared to the 23 basis points in the baseline case. Similarly, CDS spreads move by roughly 0.01 basis points in response to the same shock.[33]

Next, we examine how the model performs under a symmetric calibration of the covariance parameter ("Model 2", $\sigma_{xc\pi}(s_1) = 6 \times 10^{-4}, \sigma_{xc\pi}(s_2) = -6 \times 10^{-4}$). Under this configuration, the model generates a greater absolute bond-stock correlation in the bad regime versus the good regime, thus determining an unconditional bias toward the bad regime. This result tells us that some asymmetry in $\sigma_{xc\pi}$ (biased towards the good regime) is needed to capture the post 2000 patterns.

Finally, we focus on the role of the growth-related long-run risk parameter, Π_{cc}. Intuitively, if expected inflation shocks are embedded into x_c in a more long-lived manner, they will matter more for asset prices. Starting from the Baseline model, where $\Pi_{cc} = 0.95$, we lower this parameter to 0.85 and examine the model's performance in the final column of the table. We observe that the annualized risk premium reduces from 88 to 37 basis points, an outcome consistent with the traditional long-run risk mechanism. More interestingly, we see that the magnitudes of the regime-specific stock bond correlations, CDS sensitivities, and equity sensitivities all shrink, suggesting a lower volatility of these quantities overall.

We can more directly see this result in Figure 3, which conveys the model-implied return correlation under a lower persistence of the expected growth component ($\Pi_{cc} = 0.85$, dashed red line). Similar to the baseline case, as $\sigma_{xc\pi}$ increases, the return correlation reduces. However, the bond-stock correlation is much less sensitive to movements in the covariance term. Because expected inflation shocks are embedded for a shorter duration of time, a movement in the covariance parameter governing the expected inflation shock has less impact on the correlation of assets that embed long-term cash flows. Due to a similar logic, the

[32] Similar to the exercise in Figure 3, when changing parameters that are related to the persistence or volatility parameters of X, we make sure that the unconditional moments of X are held fixed.

[33] Any small discrepancies of Model 1 statistics across regime are due to small sample error in simulation.

magnitude of the equity return and CDS responsiveness shrink in absolute size as well. It is also worth noting that a lower persistence of the expected growth component makes it more challenging for the model to generate a negative bond-stock correlation, which is a robust feature in the data.

5 Conclusion

We shed light on the time-varying credit market sensitivity to changes in expected inflation. Using inflation swap data and high-frequency responses to macroeconomic announcements, we find that this sensitivity depends on the prevailing market perception of the inflation-growth relationship. The bond-stock return correlation emerges as a high-frequency proxy for this inflation-growth relationship and helps explain variations in market reactions. In times of market-perceived "good inflation," when inflation news is positively correlated with real economic growth, changes in expected inflation substantially reduce spreads. Meanwhile, in times of "bad inflation," the effects are reversed. A decomposition of credit spreads suggests that changes in risk premia is the channel through which inflation expectations affect spreads. A long-run risks framework provides a parsimonious economic mechanism that explains these dynamics and highlights the key role played by the nominal-real covariance.

These findings have potential implications for policy communication and financial market monitoring. In particular, the same inflation-related news may generate different asset price responses depending on broader market narratives about growth. Policymakers may benefit from recognizing this state dependence when assessing the potential effect of policy signals on financial conditions.

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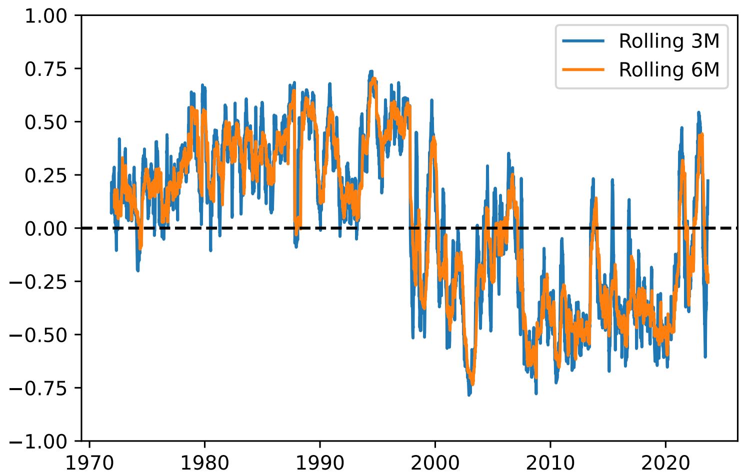
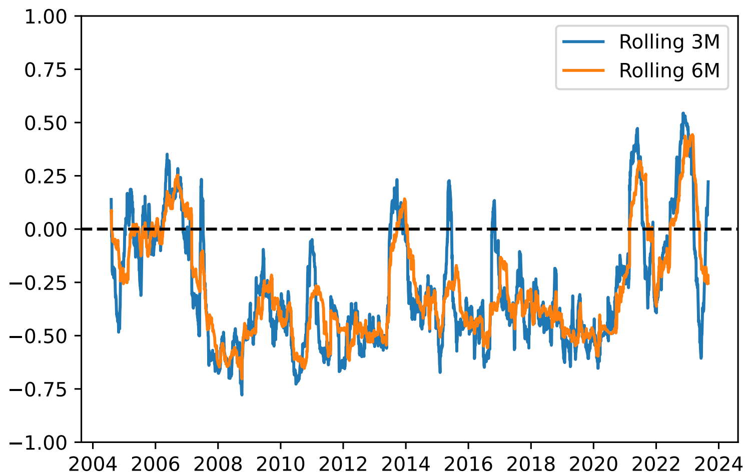
Figure 1: **Bond-Stock Return Correlation Over Time**

(a) Full Sample (Post 1970)



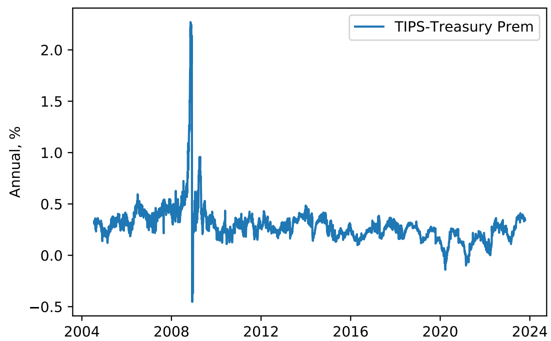
(b) Swap Sample (Post 2004)



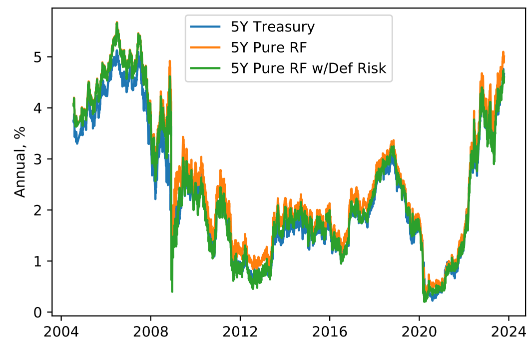
The top figure presents a time series plot of the rolling 3-month (blue) and 6-month (orange) correlation between the daily bond (5-yearU.S. Treasury) and stock market returns. The bottom figure displays the same measures over the period where inflation swaps are available (July 2004 and onwards).

Figure 2: **Bond-Stock Correlation and the Convenience Yield**

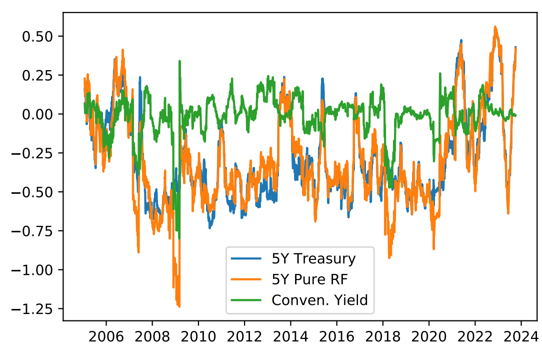
(a) Convenience Yield



(b) Frictionless Risk-Free Rates

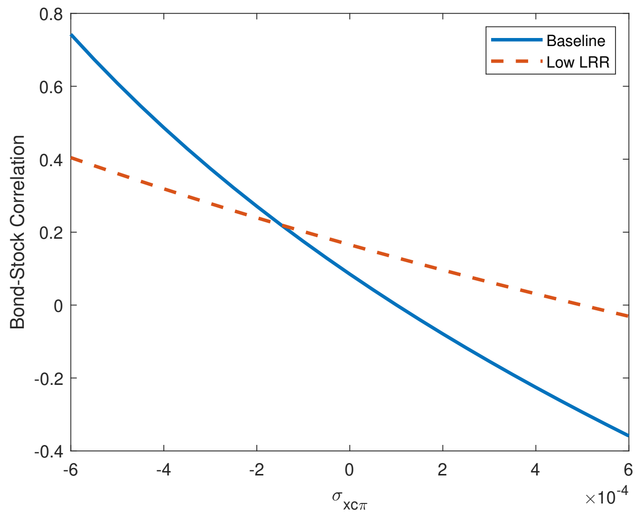


(c) Bond-Stock Corr Decomposition



The top left panel shows the time series of the convenience yield, the difference between the 5-year TIPS plus swap rate and the nominal yield. The top right panel displays the 5-year nominal yield, both in raw form and adjusted for the convenience yield and default risk. The bottom panel presents the bond–stock correlations both raw and adjusted, and the correlation between the convenience yield and stocks. See text for more detail.

Figure 3: **Model-Implied Bond-Stock Correlation and the Inflation-Growth Covariance**



This figure shows the model-implied bond-stock correlation based on simulated nominal stock and 5-year bond return data. The y-axis shows the correlation while the x-axis denotes the covariance parameter ($\sigma_{xc\pi}$). The blue line represents the bond-stock correlation across different values of $\sigma_{xc\pi}$, fixing other baseline parameters and the overall volatility of the expected growth component. The dashed red line conveys the model-implied return correlation under a lower persistence of the expected growth component ($\Pi_{cc} = 0.85$). See main text for more details.

Table 1: **Time-Varying Inflation Sensitivity of Credit Spreads**

	(1)	(2)	(3)
$\Delta\pi^{swap,5Y}$	-0.90***	-0.81***	-0.79***
	(-5.19)	(-5.27)	(-5.27)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		-0.03	
		(-0.38)	
$\tilde{\rho}_{-1}^{bond-mkt,6M}$			-0.12
			(-1.57)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.61***	
		(5.05)	
$\tilde{\rho}_{-1}^{bond-mkt,6M} \times \Delta\pi^{swap,5Y}$			0.52***
			(4.48)
$s_{i,-1}$	0.18***	0.18***	0.18***
	(3.12)	(3.21)	(3.14)
Correlation Horizon	–	3M	6M
Firm FE	Y	Y	Y
Clustering		Firm-Time	
Obs	418,777	410,129	410,129
$Adj.R^2$	0.019	0.024	0.023

This table reports the sensitivities of credit markets to movements in inflation expectations. Column (1) reports the unconditional response of CDS spreads where inflation expectations are measured by inflation swaps. Column (2) reports results where inflation swap movements are interacted with the bond-stock correlation estimated using the 3-month rolling correlation, while column (3) uses the 6-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS spread the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level.* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 2: **Risk Premia Effects and the Inflation-Growth Correlation**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta\pi^{swap,5Y}$	-0.89***	-0.27***	-0.58***	-0.82***	-0.25***	-0.53***	-0.79***	-0.25***	-0.51***
	(-5.16)	(-3.15)	(-3.89)	(-5.28)	(-3.07)	(-3.97)	(-5.24)	(-3.14)	(-3.93)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$				-0.06	-0.02	-0.04			
				(-0.85)	(-0.67)	(-0.63)			
$\tilde{\rho}_{-1}^{bond-mkt,6M}$							-0.15**	-0.03	-0.12*
							(-1.97)	(-0.98)	(-1.90)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$				0.63***	0.16**	0.44***			
				(5.15)	(2.48)	(4.16)			
$\tilde{\rho}_{-1}^{bond-mkt,6M} \times \Delta\pi^{swap,5Y}$							0.54***	0.13**	0.38***
							(4.56)	(2.01)	(3.85)
$s_{i,-1}$	0.10	0.05	0.00	0.11	0.05	0.01	0.10	0.05	-0.00
	(1.13)	(1.37)	(0.02)	(1.21)	(1.35)	(0.08)	(1.12)	(1.33)	(-0.01)
$ExpLoss_{i,-1}$	0.32***	-0.18***	0.54***	0.31***	-0.18***	0.53***	0.32***	-0.18***	0.54***
	(3.38)	(-3.22)	(5.18)	(3.26)	(-3.22)	(5.13)	(3.36)	(-3.19)	(5.25)
Dependent Variable	Δs_i (b.p.)	ΔEL_i	ΔRP_i	Δs_i (b.p.)	ΔEL_i	ΔRP_i	Δs_i (b.p.)	ΔEL_i	ΔRP_i
Correlation Horizon		–			3M			6M	
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y	Y
Clustering		Firm-Time			Firm-Time			Firm-Time	
Obs	204,172	204,150	204,148	200,303	200,281	200,279	200,303	200,281	200,279
$Adj.R^2$	0.020	0.008	0.011	0.026	0.010	0.013	0.025	0.009	0.013

This table reports the sensitivity of CDS spreads, expected losses, and credit risk premia to movements in inflation expectations. Columns (1) - (3) report unconditional results. Columns (4) - (6) report time-varying sensitvities where inflation expectation movements are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. Columns (7) - (9) report analogous results where inflation expectation movements are interacted with the inflation swap-stock correlation estimated using 6-month rolling correlation. Within each panel, from left to right, columns focus on movements in CDS spreads overall, the expected loss component, and credit risk premia. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS rate and expected loss the day before the macroeconomic announcement and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 3: **Time Varying Inflation Sensitivity Across Risk Groups**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta\pi^{swap,5Y}$	-0.82***	-0.20***	-0.77***	-2.21***	-0.53***	-0.17***	-0.59***	-1.21***
	(-5.29)	(-4.54)	(-5.38)	(-4.74)	(-3.98)	(-4.14)	(-4.17)	(-3.33)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$	-0.05	-0.02	0.01	-0.11	-0.02	-0.01	0.04	-0.02
	(-0.69)	(-0.97)	(0.17)	(-0.41)	(-0.28)	(-0.74)	(0.64)	(-0.08)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$	0.63***	0.16***	0.54***	1.74***	0.44***	0.14***	0.40***	1.06***
	(5.15)	(4.55)	(4.78)	(4.81)	(4.16)	(4.22)	(3.61)	(3.78)
$s_{i,-1}$	0.20**	0.11	0.62	0.24***	0.15**	0.09	0.61	0.19***
	(2.51)	(0.41)	(1.40)	(2.98)	(2.28)	(0.33)	(1.44)	(2.76)
Dependent Variable	Δs_i				ΔRP_i			
Which Risk Group	–	1	3	5	–	1	3	5
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
Clustering	Firm-Time				Firm-Time			
Obs	200,279	41,610	46,006	30,322	200,279	41,610	46,006	30,322
$Adj.R^2$	0.025	0.070	0.069	0.034	0.011	0.052	0.038	0.010

This table reports time-varying sensitivities of credit spreads and credit risk premia to inflation expectation movements, across different risk groups. Firms are sorted into CDS risk quintiles based on 5-year CDS spreads the day prior to macroeconomic announcements. We interact the inflation expectation movements with the bond-stock correlation estimated using the 3-month rolling correlation. Correlation measures are standardized such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. Columns (1) - (4) focus on movements in CDS spreads overall, while columns (5) - (8) on credit risk premia. Within each panel, the left most column reports the unconditional result, and the right three columns focus on risk groups 1, 3, and 5. In all regressions, we include the CDS spread on the day before macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 4: **Macro Surprises and Daily Inflation Expectation Movements**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	All Announcements				*Price-Based Announcements*			
$\Delta\pi^{swap,5Y}$	-0.90***	-0.81***			-0.95***	-0.96***		
	(-5.19)	(-5.27)			(-3.82)	(-4.14)		
$\Delta\pi^{surp,5Y}$			-0.16	-0.23*			-0.12	-0.20
			(-1.42)	(-1.93)			(-0.92)	(-1.32)
$\Delta\pi^{resid,5Y}$			-0.89***	-0.79***			-0.97***	-0.95***
			(-5.38)	(-5.46)			(-4.04)	(-4.19)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.61***				0.75***		
		(5.05)				(4.33)		
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{surp,5Y}$				0.28***				0.30***
				(3.68)				(3.21)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{resid,5Y}$				0.53***				0.65***
				(4.59)				(3.81)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		-0.03		-0.04		0.04		-0.05
		(-0.38)		(-0.51)		(0.46)		(-0.52)
$s_{i,-1}$	0.18***	0.18***	0.18***	0.18***	0.24***	0.24***	0.24***	0.24***
	(3.12)	(3.21)	(3.13)	(3.20)	(2.96)	(3.00)	(2.97)	(3.00)
Firm FE	Y		Y		Y		Y	
Clustering	Firm-Time		Firm-Time		Firm-Time		Firm-Time	
Obs	418,777	410,129	418,777	410,129	250,980	247,215	250,980	247,215
$Adj.R^2$	0.019	0.024	0.019	0.024	0.023	0.030	0.024	0.031

This table reports time-varying sensitivities of credit markets to inflation expectation movements, focusing on the macro surprise and residual component of inflation expectations. Daily movements in inflation swaps are decomposed into surprise and residual components based on a regression procedure. We then interact each of these components with the bond-stock correlation estimated using the 3-month rolling correlation. Correlation measures are standardized such that the interaction coefficient indicates the additional sensitivity to changes in the surprise or residual component of inflation swaps when the correlation is one standard deviation higher. Columns (1) - (4) focus on a decomposition using all macro surprises, while columns (5) - (8) only account for price-based surprises (CPI, PPI). In all regressions, we include the CDS spread on the day before macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 5: **Macro Surprises and Intraday Swap Prices**

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\pi^{swap,5Y}$	-1.00***	-0.85***				
	(-5.41)	(-5.12)				
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		-0.02		-0.04		-0.05
		(-0.28)		(-0.39)		(-0.59)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.59***				
		(4.34)				
$\Delta\pi^{idswap,5Y}$			-0.22	-0.28*		
			(-1.55)	(-1.79)		
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{idswap,5Y}$				0.37***		
				(2.77)		
$\Delta\pi^{surp,5Y}$					-0.12	-0.20
					(-0.89)	(-1.31)
$\Delta\pi^{latent,5Y}$					-0.34***	-0.39***
					(-2.64)	(-2.76)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{surp,5Y}$						0.23***
						(2.64)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{latent,5Y}$						0.33**
						(2.58)
$s_{i,-1}$	0.17***	0.17***	0.17**	0.18***	0.17***	0.18***
	(2.67)	(2.75)	(2.58)	(2.67)	(2.62)	(2.71)
Firm FE	Y		Y		Y	
Clustering	Firm-Time		Firm-Time		Firm-Time	
Obs	358,035	350,067	358,035	350,067	358,035	350,067
$Adj.R^2$	0.024	0.028	0.011	0.012	0.012	0.015

This table reports sensitivities of credit markets to intraday inflation expectation movements. The first two columns report results using daily inflation swaps within the same limited sample where intraday swaps are available. Column (3) examines the unconditional sensitivity of credit spreads with respect to the intraday swap change, while column (4) reports time-varying results using the same intraday swap change interacted with the 3-month bond-stock return correlation. Column (5) examines the unconditional sensitivity of credit spreads with respect to the surprise and latent components of intraday inflation swap movements. Finally, column (6) reports time-varying results using the above surprise and latent components. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS spread on the day before macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 6: **Inflation Expectations vs. Risk Premia Effects**

	(1)	(2)	(3)	(4)
$\Delta\pi^{ExpInfl}$	-0.65***			
	(-4.70)			
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{ExpInfl}$	0.41***			
	(4.58)			
$\Delta\pi^{InflRP}$		-0.47***		
		(-3.45)		
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{InflRP}$		0.34***		
		(3.48)		
$\Delta\pi^{\widetilde{ExpInfl}}$			-0.94***	
			(-6.94)	
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{\widetilde{ExpInfl}}$			0.55***	
			(5.05)	
$\Delta\pi^{\widetilde{InflRP}}$				-0.49***
				(-3.92)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{\widetilde{InflRP}}$				0.41***
				(4.94)
Decomposition Methodology	DKW		PCA	
Firm FE	Y	Y	Y	Y
Clustering	Firm-Time		Firm-Time	
Obs	410,129	410,129	403,873	403,873
Adj.R^2	0.016	0.012	0.026	0.014

This table reports the time-varying effects of daily inflation expectation and inflation risk premium movements. In the first two columns, the decomposition of inflation compensation is based on D'Amico et al. (2018), while the right two columns use a PCA procedure detailed in the Appendix. All columns report results where the inflation measures are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation measures when the correlation is one standard deviation higher. In all regressions, we include the CDS spread the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 7: **Testing Macro Measures of the Nominal-Real Covariance**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\Delta\pi^{swap,5Y}$	-0.81***	-1.01***	-0.93***	-0.78***	-0.89***	-0.83***	-0.76***
	(-5.27)	(-5.92)	(-5.35)	(-5.56)	(-6.06)	(-5.32)	(-5.59)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$	0.61***				0.53***	0.58***	0.60***
	(5.05)				(3.95)	(4.51)	(4.92)
$\widetilde{NRC}_{-1}^{EW} \times \Delta\pi^{swap,5Y}$		-0.41***			-0.20		
		(-3.59)			(-1.53)		
$\widetilde{NRC}_{-1}^{RW} \times \Delta\pi^{swap,5Y}$			-0.33***			-0.06	
			(-3.83)			(-0.63)	
$\widetilde{TCU}_{-1} \times \Delta\pi^{swap,5Y}$				0.17			0.09
				(1.34)			(0.77)
Firm FE	Y	Y	Y	Y	Y	Y	Y
Clustering		Firm-Time				Firm-Time	
Obs	410,129	418,777	418,777	418,777	410,129	410,129	410,129
$Adj.R^2$	0.024	0.021	0.021	0.021	0.024	0.024	0.025

This table reports time-varying sensitivities of credit markets to inflation expectation movements, using alternative measures of the inflation-growth covariance. Column (1) reports results using our baseline bond-stock correlation measure while column (2) reports results with an expanding window nominal real covariance measure. Column (3) uses a 60-month rolling window version of the same covariance while column (4) reports results using an adjusted version of capacity utilization. In columns (5) - (7) we run a horse race between the bond-stock correlation and alternative measures. In all regressions, we include the CDS spread the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 8: **Time-Varying Inflation Sensitivity and the Convenience Yield**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta\pi^{swap,5Y}$	-0.81***	-0.78***	-0.78***	-0.83***	-0.83***	-0.89***	-0.95***	-0.97***
	(-5.27)	(-5.14)	(-5.14)	(-5.30)	(-5.26)	(-6.22)	(-6.39)	(-5.96)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$	0.60***					0.62***		
	(5.00)					(5.47)		
$\tilde{\rho}_{-1}^{bond^{*ND}-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.39***		0.64***			0.70***	0.70***
		(3.83)		(4.79)			(5.26)	(5.56)
$\tilde{\rho}_{-1}^{bond^{*D}-mkt,3M} \times \Delta\pi^{swap,5Y}$			0.40***		0.62***			
			(3.91)		(4.89)			
$\tilde{\rho}_{-1}^{conyld-mkt,3M} \times \Delta\pi^{swap,5Y}$				-0.25**	-0.21**			0.04
				(-2.46)	(-2.15)			(0.28)
Which Sample		Full Sample		Full Sample			Non-GFC	
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
SE Clustering		Firm-Time		Firm-Time			Firm-Time	
Obs	409,903	409,903	409,903	409,903	409,903	374,600	374,600	374,600
Adj.R^2	0.024	0.023	0.022	0.025	0.025	0.017	0.018	0.018

This table reports time-varying sensitivities of credit markets to inflation expectation movements, using measures of the inflation-growth covariance that control for the convenience yield. Column (1) reports results using our baseline bond-stock correlation measure, while columns (2) and (3) report results with alternative frictionless bond-stock correlations that remove components related to the convenience yield and Treasury default risk, respectively. Columns (4) - (5) simultaneously test the frictionless component of the bond-stock correlation versus the convenience yield component, while columns (6) - (8) reports results outside the Global Financial Crisis. In all regressions, we include the CDS spread the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table 9: **Model Performance and Comparative Statics**

	Baseline	Model 1 ($\sigma_{xc\pi} = 0$)	Model 2 (Symmetric $\sigma_{xc\pi}$)	Model 3 ($\Pi_{cc} = .85$)
$E\left[pc_t\right]$	7.607	7.311	7.312	8.521
$E\left[r_{ct}\right]$	2.011	1.971	2.001	1.732
$E\left[r_{ct}^{\$}\right]$	5.538	5.498	5.528	5.259
$E\left[r_{ft}^{\$}\right]$	4.629	4.641	4.653	4.89
$E\left[r_{ct} - r_{ft}\right]$	0.908	0.857	0.875	0.368
$E\left[r_{ft}^{5Y,\$}\right]$	3.466	4.284	4.273	4.499
$E\left[s_t^{5Y}\right]$	1.337	1.332	1.326	1.285
$\sigma\left[\Delta s_t^{5Y}\right]$ (b.p.)	5.371	5.095	5.009	4.611
$\rho(r_{ct}^{\$}, r_{ft}^{5Y,\$})$	-0.148	0.085	0.073	0.09
$\rho(r_{ct}^{\$}, r_{ft}^{5Y,\$})$ – Regime 1	-0.451	0.084	-0.289	-0.079
$\rho(r_{ct}^{\$}, r_{ft}^{5Y,\$})$ – Regime 2	0.284	0.086	0.501	0.28
$\beta(\Delta s_t^{5Y} \sim \Delta x_{\pi t})$ (b.p.)	-1.603	-0.005	-0.017	-0.817
$\beta(\Delta s_t^{5Y} \sim \Delta x_{\pi t})$ – Regime 1	-6.265	0.042	-4.673	-3.242
$\beta(\Delta s_t^{5Y} \sim \Delta x_{\pi t})$ – Regime 2	3.073	-0.052	4.641	1.612
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$	0.231	-0.009	-0.006	0.072
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$ – Regime 1	0.933	-0.015	0.692	0.306
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$ – Regime 2	-0.475	-0.003	-0.705	-0.162

This table reports model moments under different parameter sets. The first column reports the endogenous model asset prices, under the baseline calibration described in the text. Model 1 is a model where the covariance channel is non-existent in both regimes ($\sigma_{xc\pi} = 0$). Model 2 sets the covariance parameter to a symmetric value across regimes ($\sigma_{xc\pi}(s_1) = 6 \times 10^{-4}$ and $\sigma_{xc\pi}(s_2) = -6 \times 10^{-4}$). Model 3 sets the long-run risk parameter (Π_{cc}) to 0.85, which is less than the baseline parameter value of $\Pi_{cc} = 0.95$.

Appendix

A Robustness and Extensions

In this Appendix, we highlight additional robustness exercises and extensions. First, we provide more specific details regarding the intraday swap analysis and PCA decomposition of inflation swaps into expectations and risk premia components. Then, we highlight other exercises supporting our main analysis, which include: various extensions of our CDS-based results to equity markets, studying the effects of inflation swap maturity, testing alternative inflation expectations measures based on TIPS breakeven rates, examining the effects of CDS liquidity, using a longer equity sample to identify sign switches in inflation sensitivities, and replacing the bond-stock correlation measure with an alternative measure based on inflation swaps.

Intraday analysis. In Section 3.2.1, we presented results based on intraday swap movements on announcement days. Appendix Table A6 details the macroeconomic announcements of interest, which include 622 announcements released monthly or quarterly at 8:30 AM ET. We also provide the number of announcements and the standard deviation of their surprises.

We use these macroeconomic announcements to examine whether swap residuals display heteroskedasticity across announcement and non-announcement days. This result is key to use the methodology of Gürkaynak et al. (2020). To do this, we compute the residual component of intraday swap movements on announcement days by regressing these movements on macroeconomic surprises. These regression results are provided in Table A7. We then compare the variance of these residuals to the variance of intraday swap movements on non-announcement days. Appendix Figure A1 displays the variance specific to different maturities and the statistical significance of the differences.

After establishing the presence of heteroskedasticity, we follow the Gürkaynak et al. (2020) methodology to identify a latent factor that is orthogonal to macroeconomic news surprises. This is done using a one-step estimator via the Kalman filter. Appendix Table A8 presents the results of this latent factor estimation from intraday swaps, showing that the latent factor is significantly related to intraday swap movements and has strong explanatory power across all horizons.

Expected Inflation and Inflation Risk Premia in Inflation Swaps. In our empirical analysis, we use daily changes in inflation swaps as a proxy for changes in expected inflation.

Because inflation swaps reflect inflation expectations under the risk-neutral measure, they are an imperfect proxy due to the presence of risk premia. In Section 3.2.2, we provide additional analysis that suggests our baseline results significantly arise from expected inflation. In addition to testing the expected inflation measure put forth by D'Amico et al. (2018), we complement their measure by constructing our own measure based on a principal component analysis. Below, we provide more details regarding this PCA.

To help identify the expected inflation component we also consider real bond yields, which have been shown to correlate negatively with expected inflation measures (e.g., Pennacchi (1991), Kandel et al. (1996), Ang et al. (2008)). Using daily changes in inflation swaps, treasury yields, and inflation-adjusted treasury yields, we extract an inflation expectation and a risk premium component, using a principal component analysis.[34] The top panel of Appendix Table A9 reports the correlations between the three variables described above and the first two principal components, which we label as the Risk Component and Expected Inflation Component. These two components explain about 90 percent of the total variation across the former three variables. In sample, we find that the risk component is positively associated with all the variables. Meanwhile, the expected inflation component is strongly positively associated with changes in both inflation compensation measures, and negatively associated with changes in inflation-adjusted Treasury yields (TIPS).[35]

The middle panel of Table A9 reports the regression of the daily changes in inflation swaps (columns (1) and (2)), break-even inflation (columns (3) and (4)), and inflation adjusted treasury yields (columns (5) and (6)) on the standardized principal components. Two facts are worth discussing. First, all three variables reflects movements in risk premia, thus validating the claim that changes in inflation swaps and breakeven inflation are only proxies for changes in market participants' expected inflation. However, the expectation component explains the great majority of daily variance in inflation swaps and breakeven inflation. Second, a standard deviation change in the expected component generates positive changes in inflation swaps and breakeven inflation that are similar in magnitude, while the change in real yields is similar in absolute magnitude but negative.

In the bottom panel, we run a battery of validation exercises for our two measures. In the first two columns, we compare daily movements in our risk component with daily changes in the Kim and Wright (2005) term premium and D'Amico et al. (2018) inflation risk premium measure, respectively. The results are comforting in that our risk premium

[34] All the variables are traded securities (hence model-free) and refer to five-year horizon. The sample starts in November 2004 and ends in October 2023.

[35] While term premia estimates might provide additional identification for the risk component, we excluded them from the PCA as they are model-dependent. That said, adding a measure of term premium (e.g., Kim and Wright (2005) or Adrian, Crump, and Moench (2013)) produces virtually the same results.

component not only is positively associated with changes in both of these measures, but also explains the majority of their daily variation. In columns (3) to (8), we assess the ability of the expectation component to predict realized CPI inflation one- and five-year ahead. To this end, we calculate the cumulative sum of the expectation component and consider its average monthly value. Then we compare the standardized version of the latter value with realized inflation at different horizons. We additionally perform the same exercise replacing our measure with the average monthly value of inflation swaps and inflation expectations reported in the University of Michigan Survey of Consumers. Here as well, the validation results are comforting. Our measure of expected inflation is the only one that correlates positively with subsequent realized inflation and that generates a reasonable R^2 value.

To conclude, we provide a visual inspection of our two components. In the top panel of Appendix Figure A2, we compare the one year-ahead realized CPI inflation (right axis) with our measure of inflation expectation (left axis). To generate the latter measure, we cumulate the expectation component and calculate its monthly average. As we can see, our measure struggles to predict one-year ahead realized inflation in some cases (for example around the Global Financial Crisis), but it captures inflation dynamics particularly well in other cases, like the 2014-2016 period and, more recently, the inflationary episode that started in 2021. In the bottom panel of Appendix Figure A2, we compare the cumulative value of our risk component with the D'Amico et al. (2018) inflation risk premium. Both quantities are at a daily level and standardized. As we can see, our risk premium component is broadly consistent with movements in inflation risk premia obtained using a no-arbitrage asset pricing framework.

Inflation Cyclicality in Equity Markets. While the focus of the paper is squarely related to credit markets, it is possible that inflation cyclicality plays a key role in the pricing of inflation expectations in equity markets. Here, we confirm that our baseline results hold using firm-level excess equity returns. While still significant in the number of firms, our equity-based tests involve a fewer number than that of the CDS sample, as our equity sample is directly matched with CDS data. Some of the results below are qualitatively similar to those in Boons et al. (2020), where the authors study a time-varying inflation risk premium that is related to the nominal-real covariance.

In Table A2, column (1) shows that equity returns, on average, react positively with respect to expected inflation news. In terms of time variation, column (2) shows that equity returns react more positively to expected inflation news, when the bond stock correlation is lower. When the correlation is 1σ lower, an increase in expected inflation, leads to a 22 b.p. higher excess equity return. Results are robust to using the slower moving 6M bond-stock correlation measure (column (3)). These results are conceptually and quantitatively in line with our CDS-based results. Expected inflation news is good for asset prices as a whole, over our post 2004 sample period, and it is only amplified when the bond-stock correlation conveys greater inflation pro-cylicality.

Systematic vs. Idiosyncratic Risk in Equity Markets. Following the earlier results on the role of inflation in credit risk premia, it is natural to ask whether equity returns also exhibit similar properties. While useful measures of equity risk premia do exist (e.g., Martin (2016), Martin and Wagner (2019)), they rely on options market data corresponding to different horizons and only characterize the lower bound of equity risk premia.[36] To circumvent these issues, we take a different approach and examine whether the equity-based inflation results in Table A2, occur through the systematic component of stock returns. To the extent that movements in expected inflation only affect returns systematically, this would provide potential evidence that equity risk premia drive our results, as the stochastic discount factor that helps determine risk premia is likely related to the same aggregate factors that determine equity returns.

Similar to Savor and Wilson (2014), we compute the systematic component of returns on a dynamic basis. For firm-level returns occurring prior to announcement date t, we conduct time series regressions in a rolling fashion:

$$\begin{aligned} R_{ik} - R_{fk} &= \alpha_{i,t-1} + \beta'_{i,t-1} X_k + \eta_{i,k} \text{ for } k = t-66, \ldots, t-1 \\ (R_{it} - R_{ft})^{sys} &= \hat{\beta}'_{i,t-1} X_t, \quad (R_{it} - R_{ft})^{idio} = (R_{it} - R_{ft}) - (R_{it} - R_{ft})^{sys}, \end{aligned}$$

where we use the past 66 days of returns to measure betas on a set of aggregate factors X. After doing so, we use the lagged betas to build the systematic and idiosyncratic components of returns: $(R_{it} - R_{ft})^{sys}$ and $(R_{it} - R_{ft})^{idio}$. The reason we use lagged betas is to ensure that systematic risk exposures are not significantly influenced by return movements on announcement days.

We substitute each of these components into our main panel specification and examine the performance in Table A3. Columns (2) and (3) provide results with respect to systematic and idiosyncratic excess returns, respectively, where we measure these components using the market factor ($X_t = [R_{mt} - R_{ft}]$). In comparison to column (1), we see that the average

[36] In unreported results, we construct the $SVIX$ and risk premia measures as given in Martin and Wagner (2019), at different option maturities. Risk premia decline on average following an increase in expected inflation, and furthermore they decline more when the bond-stock correlation is lower. However, the results are significantly affected by the financial crisis period, and the different horizons make interpretation challenging.

and time-varying sensitivities to expected inflation are completely driven by the systematic component. Meanwhile, the idiosyncratic return coefficients on $\Delta\pi^{ExpInfl}$ and its interaction effect with the bond-stock correlation are close to zero and insignificant. In columns (4) and (5), we repeat these tests but use the standard Fama and French (1993) factors related to market, size and value. Again, we find that systematic returns are the largest drivers of the average and time-varying inflation sensitivities.

Although there are certainly more elaborate sets of risk factors to test, we show here that even using a very simple design (one- or three-factor model) leads to a very robust finding: movements in expected inflation are mostly related to systematic risk. Including additional risk factors would only make these results stronger, as the systematic risk component would span a wider set of aggregate risks.

Additional Equity-Based Results. We extend our equity analysis to examine the heterogeneity in time varying inflation sensitivity. In Table A4, we sort firms into quintiles based on their CDS spread on the day prior to the macro announcement. We then examine the expected inflation sensitivities within each group of firms. Comparing columns (2) and (4), we find that the riskiest firms have a significantly larger average and time-varying coefficient. Similar to the CDS-based results on heterogeneity, riskier firms display wider swings in their inflation sensitivity, based on the bond-stock correlation.

Similar to the exercise in the main text, examining the role of macro surprises and residual inflation expectations for the inflation sensitivities of CDS spread movements, we also study which component plays an important role in equities. Table A5 provides the results. Using an identical decomposition to earlier, we show in columns (3) and (4), that the residual component of inflation expectations plays an overwhelming role for the average effect. For the time-varying effect the residual is significant, in addition to the macro surprise component.

Inflation Sensitivities by Swap Horizon. Inflation swaps trade at multiple maturities which might suggest that our results vary with respect to swap maturity. To explore this idea, we replace our main inflation measure (change in five-year inflation swap rates) with movements in the one- and ten-year inflation swap rate. Results are reported in Table A10. As shown in column (1), the average and time-varying effects of swap movements are greatly attenuated for CDS spreads when examining the one-year horizon. Meanwhile the ten-year horizon (column (3)) is very similar to our baseline (middle column). Results are even more stark with respect to equity returns where the one-year swap is insignificant altogether (column (4)). Our results are intuitive from the standpoint that equities and corporate credit spreads reflect the health of long-duration assets subject to longer term fundamentals. To this point, Bahaj et al. (2023) show using regulatory data on inflation swap trading in the U.K., that short term inflation swap prices are mostly informed by liquidity shocks, while fundamentals play a stronger role at longer maturities. The dominance of liquidity shocks at the short horizon might partially be due to market segmentation, as they show that hedge funds tend to trade in and out of short-term positions, while the longer end is more so driven by buy-and-hold pension funds looking for inflation protection.

Response to Breakeven Inflation. As we show in the right panel of Figure A3, TIPS-based inflation expectations (constant maturity five-year nominal yield minus constant maturity five-year TIPS), broadly tracks well with our swap measure. To ensure that our results are not specific to the inflation measure we have chosen, we re-conduct our main analysis using five-year breakeven inflation, $\Delta\pi^{be,5Y}$.

Appendix Table A11 shows that our main results are robust when we account for breakeven inflation. The first column shows that five-year CDS declines by 1.0 basis point, following a standard deviation movement in five-year breakeven inflation, surrounding macroeconomic announcements. Analogously, equity returns rise by 37 basis points following the movement in breakeven inflation. Columns (3) and (4) show, similar to results using inflation swap changes, that the large majority of the effect comes through risk premia effects. The final two columns suggest that the time-variation that was earlier documented also holds when looking at breakeven inflation. A more negative bond-stock correlation is associated with larger CDS declines and equity price appreciation, following an increase in inflation expectations as measured through TIPS breakeven inflation. These results confirm that our fundamental economic mechanism holds regardless of the expected inflation measure.

CDS Liquidity. The Dodd-Frank Act and additional regulations have led to greater standardization and regulation of CDS trading, reducing the size of the single-name CDS market over time (e.g., Boyarchenko et al. (2020)). Consequently, it is important to assess whether our results are affected by low liquidity in CDS markets. Appendix Table A12 shows that our main results are robust in different CDS liquidity samples. We examine the number of participating dealers for a given reference entity, as a greater number of dealers might indicate higher liquidity. We compute the cross-sectional median number of dealers on each announcement date, and we report results across different groups (greater and less than the median number of dealers). Our results strengthen when focusing on firms with a larger number of dealers on each announcement day, while results also hold for firms with a low number of dealers CDS.

Time-Varying Inflation Sensitivities Including Non-Announcement Days. In our main analysis, we focus on macroeconomic announcement days, when investors are most likely to incorporate new information into asset prices. These days exhibit heightened market activity, with inflation swap movements reflecting the arrival and pricing of economic news. However, announcement days are relatively infrequent—comprising a small fraction of our full sample—and exhibit considerably higher volatility. Specifically, the variance of daily inflation swap movements is between 2 and 3.5 times larger on announcement days compared to non-announcement days.

To evaluate whether our results are specific to high-volatility, news-driven environments or instead reflect a broader pricing mechanism, we re-estimate our main regression over the entire sample period, including both announcement and non-announcement days. Appendix Table A13 replicates the specification from Table 1, extending the analysis to all trading days from August 2004 to October 2023.

Consistent with our expectations, we find that the estimated coefficients—both for the unconditional effect of inflation expectations and for their interaction with the bond-stock correlation—are attenuated in magnitude relative to the announcement-day results. This attenuation reflects the lower variance in swap price movements and reduced information flow on non-announcement days. Nonetheless, the coefficients remain statistically significant, confirming that our main findings hold overall. These results suggest that the economic mechanism we identify is not solely dependent on macroeconomic announcements, but also manifests, albeit more weakly, on typical trading days.

Inflation-Growth Regimes over a Long Sample. As is well documented, the bond-stock return correlation significantly changed sign in the late 1990s, turning from positive to negative. Because our sample focuses on the post-2004 period, it is difficult to detect discrete sign switches in inflation responsiveness. To understand whether sign switches are a possibility, we extend back our credit sample using the aggregate, ICE BofA US Corporate Option-Adjusted Spread index, available from 1996. Additionally, we extend the equity return panel used in Table A2 back to the 1980s. Using the daily inflation measures from D'Amico et al. (2018) surrounding macroeconomic announcements, we test for sign switches with respect to inflation sensitivities.

In addition to the tests from our baseline analysis, we modify our interaction regression to include a dummy variable in addition to the standardized correlation measure:

$$\begin{aligned} \Delta ys_t &= \beta_0 + \beta_\pi \Delta\pi_t^{InfComp} + \beta_{\rho\pi}\left(\mathbb{1}_{\{\rho_{t-1}>0\}} \times \Delta\pi_t^{InfComp}\right) + \beta_X' X_{t-1} + \varepsilon_t \\ r_{it} - r_{ft} &= \beta_i + \beta_\pi \Delta\pi_t^{InfComp} + \beta_{\rho\pi}\left(\mathbb{1}_{\{\rho_{t-1}>0\}} \times \Delta\pi_t^{InfComp}\right) + \beta_X' X_{i,t-1} + \varepsilon_{it} \end{aligned} \quad \text{(A1)}$$

where in the top regression, Δys_t indicates the aggregate, daily change in yield spreads. Using the correlation measure based on risk-free bond and stock returns, we interact the inflation measure change with a dummy variable ($\mathbb{1}_{\{\rho_{t-1}>0\}}$), which indicates whether the raw correlation (non-standardized) is positive, which is interpretable as a "bad inflation" state. Breaking up the regimes in this way will also tell us whether the bad inflation regime shows statistically different behavior than a good one.

We provide results for this test in Appendix Table A14 and Appendix Table A15. We first show that the time-varying results hold in the extended sample. Columns (1) and (2), and (5) and (6) report results using the bond-stock correlation at the three-month or six-month horizon. Using either total inflation compensation or physical inflation expectations, the time-varying coefficients are similar in magnitude to the ones in the baseline sample. Next, in columns (3) and (4), and (7) and (8) we report the results accounting for correlation regimes. It is evident that the bad inflation regime displays statistically *more positive* (*more negative*) responses to inflation movements than in the good regime in credit (equity). Furthermore, in equities, the response to inflation news in the $\rho > 0$ regime is negative overall ($-0.536 + 0.341 < 0$). Both of these results validate our original hypothesis. We show that indeed in negative (positive) correlation regimes the equity sensitivity is positive (negative) and the credit reaction is muted. These results suggest the good and bad inflation pricing dynamics are present over a longer time span.

Swap-Based Correlation Measure. Our results have focused on time-variation using the bond-stock return correlation as a key statistic. In this exercise, we use an alternative measure which correlates daily changes in inflation swap prices to market returns. In Appendix Figure A5 we display a plot of this measure over time. Because movements in swap rates positively correlate with inflation risk and yield movements, it is approximately the flipped image of the original bond-stock correlation measure displayed in the bottom of Figure 1. Over the past two decades it has remained mostly positive with short periods where it turns negative.

We replace our bond-based correlation measure with a swap-based one and re-examine our main regressions. Appendix Table A16 displays these results. As shown through the CDS results (left three columns), regardless of the three-month or six-month horizon, increases in the prior swap-market correlation (more of a good inflation environment) lead to a further reduction in spreads following an expected inflation movement. Equity markets provide a qualitatively similar result. All told, using the swap-based correlation measure does not affect our results and in some cases increases the statistical significance.

B CDS Decomposition

CDS spreads at a given maturity is the annualized rate C_t, such that:

$$\Delta C_t \sum_{k=1}^{K/\Delta} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k\Delta} \left(1 - D_{t,(k-1)\Delta}\right) \right] = \sum_{k=1}^{K/\Delta} \mathbb{E}_t \left[\tilde{M}^{\$}_{t+k\Delta} \times L_{t+(k-1)\Delta,\Delta} \times D_{t+(k-1)\Delta,\Delta} \right] \tag{A2}$$

The only difference relative to Equation (12) is that we allow for losses given default to be time-varying above. By definition, the expected loss component is one where we assume risk neutrality of the SDF. Along with two other assumptions (conditional independence of recovery rates from realized default and martingale nature of recovery rates), one can transform the above equation to receive:

$$ExpLoss_t = \frac{L_t \sum_{k=1}^{K/\Delta} d_{t,k\Delta} \mathbb{E}_t \left[D_{t+(k-1)\Delta,\Delta} \right]}{\Delta \sum_{k=1}^{K/\Delta} d_{t,k\Delta} \mathbb{E}_t \left[1 - D_{t,(k-1)\Delta} \right]} \tag{A3}$$

where $ExpLoss_t$ is the expected loss component and $d_{t,k\Delta}$ is the time t discount rate of a cash flow at $t + k\Delta$. Inherent in this expression is that the decomposition is firm, time, and maturity specific.

While Berndt et al. (2018) compute $ExpLoss_{it}$ using this nonlinear functional form, we use the approximation from Palazzo and Yamarthy (2022), where the authors show that $L_t \times \mathbb{E}_t \left[D_{t+(k-1)\Delta,\Delta} \right]$, the product of loss given default and the (annualized) probability of default over the course of the CDS contract, is close in level terms and highly correlated to the fully nonlinear form that accounts for the term structure of default probabilities. Using this approximation is convenient as it is a straightforward formula requiring two pieces of data: recovery rate estimates (available from Markit) and default probability estimates (from Moody's). After obtaining $ExpLoss_{it}$, the credit risk premium is defined as the additive residual, $RiskPrem_{it} = s_{it} - ExpLoss_{it}$.

C Model Solution

C.1 Price-to-Consumption Ratio

Based on the Euler equation restriction and fundamental assumptions we can show that the price-consumption ratio takes the form:

$$pc_t = A_1' X_t + A_2(s_t) \tag{A4}$$

where A_1 is a set of loadings on expected growth and inflation and A_2 is a regime switching component. To show this we start with the Euler Equation:

$$\mathbb{E}_t\left[\exp\left(m_{t+1} + r_{c,t+1}\right)\right] = \mathbb{E}_t\left[\exp\left(\theta \log \delta - \frac{\theta}{\psi}\Delta c_{t+1} + \theta r_{c,t+1}\right)\right] = \exp(0) \tag{A5}$$

$$(\Longleftrightarrow) \exp(\theta pc_t) = \mathbb{E}_t\left[\exp\left(\theta \log \delta + (1-\gamma)\Delta c_{t+1} + \theta\kappa_0 + \theta\kappa_1 pc_{t+1}\right)\right]$$

We guess / verify the *pc* guess and simplify the right hand side:

$$\begin{aligned}
\exp(\theta pc_t) &= \mathbb{E}_t\left[\exp\left(\theta \log \delta + (1-\gamma)\Delta c_{t+1} + \theta\kappa_0 + \theta\kappa_1 pc_{t+1}\right)\right] \\
&= \mathbb{E}_t\left[\exp\left\{(1-\gamma)\sigma_c \varepsilon_{c,t+1} + \theta\kappa_1 pc_{t+1}\right\}\right] \\
&\quad \times \exp\left(\theta \log \delta + (1-\gamma)\mu_c + \theta\kappa_0 + (1-\gamma)e_1' X_t\right) \\
&= \mathbb{E}_t\left[\exp\left(\theta\kappa_1 A_1' \Sigma_t \eta_{t+1}\right)\right] \times \mathbb{E}_t\left[\exp\left(\theta\kappa_1 A_2(s_{t+1})\right)\right] \\
&\quad \times \exp\left(\frac{1}{2}(1-\gamma)^2\sigma_c^2 + \theta\kappa_1 A_1' \Pi X_t\right) \\
&\quad \times \exp\left(\theta \log \delta + (1-\gamma)\mu_c + \theta\kappa_0 + (1-\gamma)e_1' X_t\right) \\
&= \underbrace{\exp\left(\frac{1}{2}\theta^2\kappa_1^2 A_1' \Sigma_t \Sigma_t' A_1\right) \times \exp\left(\log\left\{\sum_{j=1}^{2} p_{ij} \exp\left(\theta\kappa_1 A_2(s_j)\right)\right\}\right)}_{\text{Dependent on } s_t} \\
&\quad \times \exp\left(\frac{1}{2}(1-\gamma)^2\sigma_c^2 + \theta\kappa_1 A_1' \Pi X_t\right) \\
&\quad \times \exp\left(\theta \log \delta + (1-\gamma)\mu_c + \theta\kappa_0 + (1-\gamma)e_1' X_t\right)
\end{aligned} \tag{A6}$$

Matching coefficients on X_t we receive:

$$\begin{aligned}
\theta A_1' &= (1-\gamma)e_1' + \theta\kappa_1 A_1' \Pi \\
A_1 &= (1 - \frac{1}{\psi}) \times (I - \kappa_1 \Pi')^{-1} e_1
\end{aligned} \tag{A7}$$

Matching coefficients on s_t we receive:

$$\begin{aligned}\theta A_2(s_t=i) = \theta\log\delta + (1-\gamma)\mu_c + \theta\kappa_0 + \frac{1}{2}(1-\gamma)^2\sigma_c^2 \\ + \frac{1}{2}\theta^2\kappa_1^2 A_1'\Sigma_t\Sigma_t' A_1 + \log\left\{\sum_{j=1}^{N} p_{ij}\exp\left(\theta\kappa_1 A_2(s_j)\right)\right\} \\ \text{for } i=1,\dots,N\end{aligned} \tag{A8}$$

This is a system of N equations and N unknowns that we can solve numerically.

C.2 Nominal Bond Returns

The return on an n-period zero-coupon bond return (purchase at t, sell at $t+1$) will be given by:

$$\exp\left(r_{f,t+1}^{\$,n}\right) = \frac{P_{f,t+1}^{\$,n-1}}{P_{f,t}^{\$,n}} = \exp\left(p_{f,t+1}^{\$,n-1} - p_{f,t}^{\$,n}\right) \tag{A9}$$

where $P_{f,t}^{\$,n}$ indicates the price of a risk-free bond at time t that matures at $t+n$, and its lowercase is in log terms. We can show that the log price will take the form:

$$p_{f,t}^{\$,n} = P_1^{n'} X_t + P_2^n(s_t) \tag{A10}$$

Starting with $n=1$ (one period risk-free bond), we have:

$$\begin{aligned}\exp\left(p_{ft}^{\$,1}\right) &= \mathbb{E}_t\left[\exp\left(m_{t+1}-\pi_{t+1}\right)\right] \\ &= \mathbb{E}_t\left[\exp\left(\theta\log\delta - \frac{\theta}{\psi}\Delta c_{t+1} - (1-\theta)r_{c,t+1} - \pi_{t+1}\right)\right] \\ &= \mathbb{E}_t\left[\exp\left(\theta\log\delta - (1-\theta+\frac{\theta}{\psi})\Delta c_{t+1} - (1-\theta)(\kappa_0+\kappa_1 pc_{t+1} - pc_t) - \pi_{t+1}\right)\right] \\ &= \exp\left(\theta\log\delta - (1-\theta+\frac{\theta}{\psi})(\mu_c + e_1' X_t) - (1-\theta)(\kappa_0 + \kappa_1 A_1'\Pi X_t - pc_t) - (\mu_\pi + e_2' X_t)\right) \\ &\quad\times \mathbb{E}_t\left[\exp\left((1-\theta+\frac{\theta}{\psi})\sigma_c\varepsilon_{c,t+1} - (1-\theta)\left(\kappa_1 A_1'\Sigma_t\eta_{t+1} + \kappa_1 A_2(s_{t+1})\right) - \sigma_\pi\varepsilon_{\pi,t+1}\right)\right]\end{aligned} \tag{A11}$$

Final price can be expressed as:

$$\begin{aligned}p_{ft}^{\$,1} = \theta\log\delta - (1-\theta+\frac{\theta}{\psi})\mu_c - (1-\theta)\kappa_0 - \mu_\pi + \frac{1}{2}(1-\theta+\frac{\theta}{\psi})^2\sigma_c^2 + \frac{1}{2}\sigma_\pi^2 \\ + (1-\theta)A_2(s_t) + \frac{1}{2}(1-\theta)^2\kappa_1^2 A_1'\Sigma_t\Sigma_t' A_1 + \log\left(\mathbb{E}_t\left[\exp\left((\theta-1)\kappa_1 A_2(s_{t+1})\right)\right]\right) \\ + \left[(\theta-1-\frac{\theta}{\psi})e_1' - (1-\theta)\kappa_1 A_1'\Pi + (1-\theta)A_1' - e_2'\right] X_t\end{aligned} \tag{A12}$$

where $P_1^{1'}$ is indicated by the coefficient in the brackets in the third line, and $P_2^1(s_t)$ is indicated by the top two lines.

To solve for a maturity n, assume that the statement holds for $n-1$, i.e. that there exist coefficients such that $p_{ft}^{\$,n-1} = P_1^{n-1'} X_t + P_2^{n-1}(s_t)$. Due to the zero-coupon nature of these bonds:

$$\exp\left(p_{ft}^{\$,n}\right) = \mathbb{E}_t\left[\exp\left(m_{t+1} - \pi_{t+1} + p_{f,t+1}^{\$,n-1}\right)\right] \tag{A13}$$

as the price will be the nominally discounted value of the future market value. We can further simplify:

$$\begin{aligned}\exp\left(p_{ft}^{\$,n}\right) &= \mathbb{E}_t\left[\exp\left(\theta\log\delta - (1-\theta+\frac{\theta}{\psi})\Delta c_{t+1} - (1-\theta)(\kappa_0 + \kappa_1 pc_{t+1} - pc_t) - \pi_{t+1} + P_1^{n-1'} X_{t+1} + P_2^{n-1}(s_{t+1})\right)\right] \\ &= \exp\left(\theta\log\delta - (1-\theta+\frac{\theta}{\psi})(\mu_c + e_1' X_t) - (1-\theta)(\kappa_0 + \kappa_1 A_1'\Pi X_t - pc_t) - (\mu_\pi + e_2' X_t) + P_1^{n-1'}\Pi X_t\right) \\ &\times \mathbb{E}_t\left[\exp\left((1-\theta+\frac{\theta}{\psi})\sigma_c\varepsilon_{c,t+1} - (1-\theta)\left(\kappa_1 A_1'\Sigma_t\eta_{t+1} + \kappa_1 A_2(s_{t+1})\right) + P_1^{n-1'}\Sigma_t\eta_{t+1} + P_2^{n-1}(s_{t+1}) - \sigma_\pi\varepsilon_{\pi,t+1}\right)\right]\end{aligned} \tag{A14}$$

The final price can be written as:

$$\begin{aligned}p_{ft}^{\$,n} &= \theta\log\delta - (1-\theta+\frac{\theta}{\psi})\mu_c - (1-\theta)\kappa_0 - \mu_\pi + \frac{1}{2}(1-\theta+\frac{\theta}{\psi})^2\sigma_c^2 + \frac{1}{2}\sigma_\pi^2 \\ &+ (1-\theta)A_2(s_t) + \frac{1}{2}\left(P_1^{n-1'} - (1-\theta)\kappa_1 A_1'\right)\Sigma_t\Sigma_t'\left(P_1^{n-1'} - (1-\theta)\kappa_1 A_1'\right)' \\ &+ \log\left(\mathbb{E}_t\left[\exp\left\{(\theta-1)\kappa_1 A_2(s_{t+1}) + P_2^{n-1}(s_{t+1})\right\}\right]\right) \\ &+ \left[(\theta - 1 - \frac{\theta}{\psi})e_1' - (1-\theta)\kappa_1 A_1'\Pi + (1-\theta)A_1' - e_2' + P_1^{n-1'}\Pi\right] X_t\end{aligned} \tag{A15}$$

The coefficients for $\left\{P_1^{n'}, P_2^n(s_t)\right\}$ are a function of the maturity $n-1$ coefficients. Using these one can compute nominal bond prices and corresponding bond returns.

C.3 CDS Spreads

As given in Equation (13) of the main text, we need to compute two quantities to solve the model:

$$\underbrace{\mathbb{E}_t\left[\tilde{M}_{t+k}^{\$} S_{t,t+k}\right]}_{(*)}, \quad \underbrace{\mathbb{E}_t\left[\tilde{M}_{t+k}^{\$} S_{t,t+k-1}\right]}_{(**)}$$

taking into account the nominal SDF assumptions of the model, long-run risk fundamentals, and exogenous default dynamics:

$$D_{t,1} = \begin{cases} 0 & \text{w/probability } \exp(-\lambda_t) \\ 1 & 1-\exp(-\lambda_t) \end{cases} \tag{A16}$$

$$\lambda_t = \beta_{\lambda 0}(s_t) + \beta'_{\lambda x} X_t$$

Key Analytical Result Before simplifying the expectational terms, we mention a key analytical result. Suppose we have a generic function, $f_t = f'_1 X_t + f_2(s_t)$, then we can show that there exists coefficients for $\tilde{f}_t$ such that:

$$\begin{aligned} \tilde{f}(s_t, x_t) &= \mathbb{E}_t \left[\tilde{M}^{\$}_{t+1} \times \exp\left(f'_1 X_{t+1} + f_2(s_{t+1})\right)\right] \\ &= \mathbb{E}_t \left[\exp\left(m_{t+1} - \pi_{t+1} + f'_1 X_{t+1} + f_2(s_{t+1})\right)\right] \\ &= \exp\left(\tilde{f}'_1 X_t + \tilde{f}_2(s_t)\right) \end{aligned} \tag{A17}$$

The coefficients for $\tilde{f}$ are given by:

$$\begin{aligned} \tilde{f}_2(s_t) = {} & \theta \log\delta - (1-\theta+\frac{\theta}{\psi})\mu_c - (1-\theta)\kappa_0 - \mu_\pi + \frac{1}{2}(1-\theta+\frac{\theta}{\psi})^2\sigma_c^2 + \frac{1}{2}\sigma_\pi^2 \\ & + (1-\theta)A_2(s_t) + \frac{1}{2}\left(f'_1 - (1-\theta)\kappa_1 A'_1\right)\Sigma_t\Sigma'_t\left(f'_1 - (1-\theta)\kappa_1 A'_1\right)' \\ & + \log\left(\mathbb{E}_t\left[\exp\left\{(\theta-1)\kappa_1 A_2(s_{t+1}) + f_2(s_{t+1})\right\}\right]\right) \end{aligned} \tag{A18}$$

$$\tilde{f}'_1 = \left[(\theta - 1 - \frac{\theta}{\psi})e'_1 - (1-\theta)\kappa_1 A'_1 \Pi + (1-\theta)A'_1 - e'_2 + f'_1 \Pi\right]$$

Solving for (*) We can rewrite the expression as:

$$\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k} S_{t,t+k}\right] = \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k} \Pi_{j=1}^k S_{t+j-1,1}\right] = \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\right] \tag{A19}$$

where the right most term uses the conditional independence default assumption. For $k=1$, this term simplifies to:

$$\begin{aligned} \mathbb{E}_t\left[\tilde{M}^{\$}_{t+1} S_{t,t+1}\right] = \exp(-\lambda_t) \times \mathbb{E}_t\left[\tilde{M}^{\$}_{t+1}\right] &= \exp\left(p^{\$,1}_{ft} - \beta'_{\lambda x} X_t - \beta_{\lambda 0}(s_t)\right) \\ &= \exp\left((P^1_1 - \beta_{\lambda x})' X_t + P^1_2(s_t) - \beta_{\lambda 0}(s_t)\right) \\ &= \exp\left(B^{1'}_1 X_t + B^1_2(s_t)\right) \end{aligned} \tag{A20}$$

For $k > 1$, the right most term can be simplified to:

$$\begin{aligned}
\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\right] &= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-1}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\mathbb{E}_{t+k-1}\left[\exp\left(m_{t+k}-\pi_{t+k}\right)\right]\right]\\
&= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-1}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\exp\left(p^{\$,1}_{f,t+k-1}\right)\right]\\
&= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-1}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\exp\left(P_1^{1'}X_{t+k-1}+P_2^1(s_{t+k-1})\right)\right]
\end{aligned} \tag{A21}$$

Given all terms on the RHS are at the $t+k-1$ timestep we can apply the result from earlier. Sequentially, we compute the expectation:

$$\begin{aligned}
&\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-1}\exp\left(-\sum_{j=1}^{k}\lambda_{t+j-1}\right)\exp\left(P_1^{1'}X_{t+k-1}+P_2^1(s_{t+k-1})\right)\right]=\\
&\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\mathbb{E}_{t+k-2}\left[M^{\$}_{t+k-1}\times\exp\left(P_1^{1'}X_{t+k-1}+P_2^1(s_{t+k-1})-\lambda_{t+k-1}\right)\right]\right]=\\
&\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\exp\left(\tilde{P}_1X_{t+k-2}+\tilde{P}_2(s_{t+k-2})\right)\right]=\\
&\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-3}\exp\left(-\sum_{j=1}^{k-2}\lambda_{t+j-1}\right)\mathbb{E}_{t+k-3}\left[M^{\$}_{t+k-2}\times\exp\left(\tilde{P}_1'X_{t+k-2}+\tilde{P}_2(s_{t+k-2})-\lambda_{t+k-2}\right)\right]\right]=\\
&\cdots=\exp\left(B_1^{k'}X_t+B_2^k(s_t)\right)
\end{aligned} \tag{A22}$$

where to get from the second to third line, we use the earlier result. The final expression is exponential affine in the expected growth / inflation state and the Markov state.

Solving for (**) The proof will be similar to the solution for $(*)$. We can rewrite the expression as:

$$\begin{aligned}
\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}S_{t,t+k-1}\right] &= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}\Pi_{j=1}^{k-1}S_{t+j-1,1}\right]\\
&= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\right]
\end{aligned} \tag{A23}$$

where the right most term uses the conditional independence default assumption. For $k = 1$, this term simplifies to:

$$\mathbb{E}_t\left[\tilde{M}^{\$}_{t+1}S_{t,t}\right] = \mathbb{E}_t\left[\tilde{M}^{\$}_{t+1}\right] = \exp\left(p^{\$,1}_{ft}\right) = \exp\left(C_1^{1'}X_t + C_2^1(s_t)\right) \tag{A24}$$

For $k > 1$, the right most term can be simplified to:

$$\begin{aligned}
\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\right] &= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\mathbb{E}_{t+k-2}\left[M^{\$}_{t+k-1}\times M^{\$}_{t+k}\right]\right] \\
&= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\exp\left(p^{\$,2}_{f,t+k-2}\right)\right] \\
&= \mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\exp\left(P_1^{2'}X_{t+k-2}+P_2^2(s_{t+k-2})\right)\right]
\end{aligned} \tag{A25}$$

Given all terms on the RHS are at the $t+k-2$ timestep we can apply the result from earlier. Sequentially, we compute the expectation and receive similar to earlier that:

$$\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k-2}\exp\left(-\sum_{j=1}^{k-1}\lambda_{t+j-1}\right)\exp\left(P_1^{2'}X_{t+k-2}+P_2^2(s_{t+k-2})\right)\right] = \exp\left(C_1^{k'}X_t + C_2^k(s_t)\right) \tag{A26}$$

The final expression is exponential affine in the expected growth / inflation state and the Markov state.

Overview Based on the solutions for $\left\{B_1^k, B_2^k(s_t), C_1^k, C_2^k(s_t)\right\}$ we can write the 5Y CDS as:

$$\begin{aligned}
C_t &= (1-R)\times\left(1-\frac{\sum_{k=1}^{20}\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}S_{t,t+k}\right]}{\sum_{k=1}^{20}\mathbb{E}_t\left[\tilde{M}^{\$}_{t+k}S_{t,t+k-1}\right]}\right) \\
&= (1-R)\times\left(1-\frac{\sum_{k=1}^{20}\exp\left(B_1^{k'}X_t + B_2^k(s_t)\right)}{\sum_{k=1}^{20}\exp\left(C_1^{k'}X_t + C_2^k(s_t)\right)}\right)
\end{aligned} \tag{A27}$$

D Calibration Details

As it is standard in the long-run risks literature, we numerically calibrate the model at a quarterly frequency. That said, the mechanisms we discuss hold at higher frequencies, as we show in our empirical analysis. In the calibration, we make two simplifying assumptions. First, the autoregressive matrix Π is set to be diagonal with no cross dependencies. This assumption allows for a clean interpretation of the covariance parameter as the sole source of the real-nominal interaction. Second, we fix the number of regimes to $N = 2$ so that we can speak to distinctive “good” and “bad” inflation regimes.

Appendix Table A17 lists the baseline parameter values (top panel). Some parameters are taken from the literature (e.g., $\gamma, \delta, \Pi_{cc}, \Pi_{\pi\pi}$) while others are calibrated. Putting aside the inflation-growth covariance parameter, we calibrate the fundamental parameters (those of $\Delta c, \pi$) to match, or get reasonably close to first and second moments of consumption growth and inflation, between 1968:Q4 and 2019:Q4.[37] We also match the unconditional volatilities of expected real growth and inflation, constructed using survey data from the Survey of Professional Forecasters (SPF) and the methodology in Bansal and Shaliastovich (2012).

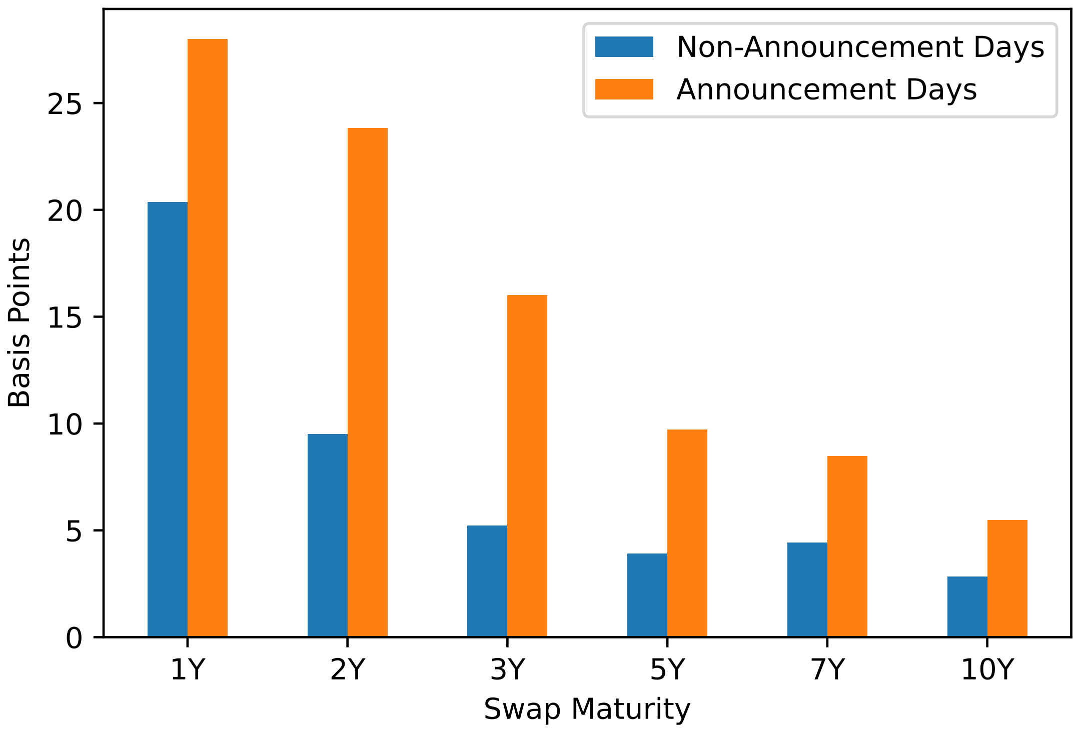
As shown in Figure 3, the "good inflation" regime with $\sigma_{xc\pi} > 0$ produces a negative stock-bond correlation, while the "bad inflation" regime produces the opposite. Because much of our data sample (post 2000s) lies in the former, we calibrate $|\sigma_{xc\pi}(s_1)| > |\sigma_{xc\pi}(s_2)|$, with $\sigma_{xc\pi}(s_1) > 0$ and $\sigma_{xc\pi}(s_2) < 0$. Hence s_1 is our good inflation regime, where orthogonal shocks to expected inflation feedback positively to expected growth. Conditional transition probabilities on the regime (p_{11}, p_{22}) are chosen to be equal, with an average regime length of 8 to 10 quarters.

In terms of credit parameters we calibrate the recovery rate (R) and default parameters $(\beta_{\lambda 0}, \beta_{\lambda x})$ which govern the hazard rate function. We set $R = 0.4$ in line with the panel average of Markit recovery rates. To simplify the model $\beta_{\lambda 0} = 0.505$ percent across both regimes to target a 2 percent annual default rate, close to the empirical average.[38] Finally, we only allow λ_t to depend on x_{ct} as default rates tend to significantly correlate with economic growth measures. We calibrate $\beta_{\lambda xc} < 0$ to generate reasonable countercyclicality of default rates and volatility of CDS spreads.

[37] We do not include data beyond 2019:Q4 to avoid the extreme volatility induced by the COVID-19 episode.

[38] Based on Moody's EDF data, the average annualized five-year default probability is roughly 1.1 percent. We calibrate average default rates a bit higher to get closer to the CDS spread level in the data.

Figure A1: **Heteroskedasticty of Intraday Swap Residuals**

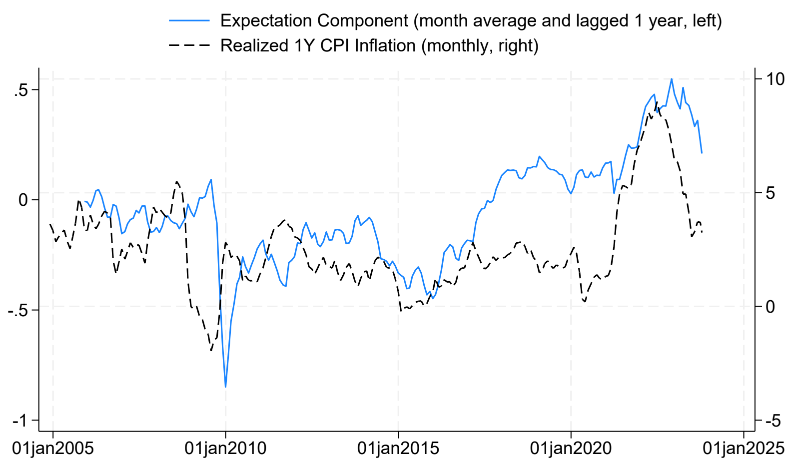


	1Y	2Y	3Y	5Y	7Y	10Y
var(η_t^A)	28.00	23.83	16.02	9.72	8.49	5.49
var(η_t^{NA})	20.37	9.50	5.23	3.90	4.44	2.84
F-test Statistic	1.37***	2.51***	3.06***	2.49***	1.91***	1.93***

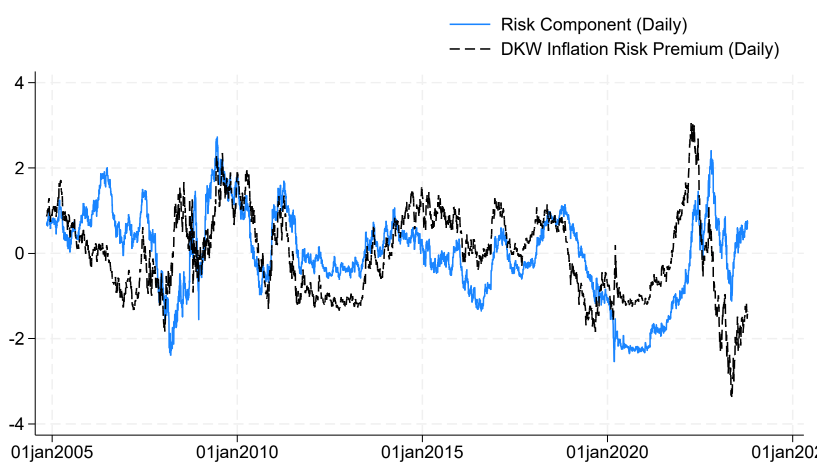
This figure display the maturity-specific variance of intraday inflation swap movements on announcement and non-announcement days. For announcement days, the variance is computed using the portion of intraday swap changes that is not related to macroeconomic surprises, via regression residuals. Meanwhile, for non-announcement days the raw swap change is used to compute the variance. Inflation swap data is collected daily from 8:15 AM to 9:15 AM ET, reflecting a 60-minute window. The table below reports the variance in basis points, and a F-test statistic regarding the significance of the difference.

Figure A2: **PCA-Derived Inflation Expectations and Risk Premium**

(a) Inflation Expectations Component



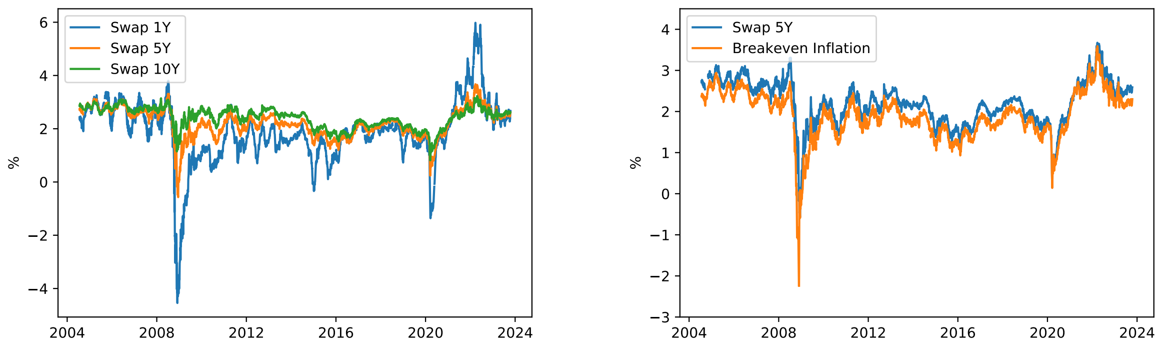
(b) Inflation Risk Premium Component



The top figure shows the monthly average of the cumulative expectation component (blue) compared to one-year-ahead realized CPI inflation (black, right axis). The bottom figure displays the cumulative value of the standardized risk component (blue) alongside the standardized inflation risk premium from D'Amico et al. (2018) (black), both computed at a daily frequency.

Figure A3: **CPI Inflation Swaps**

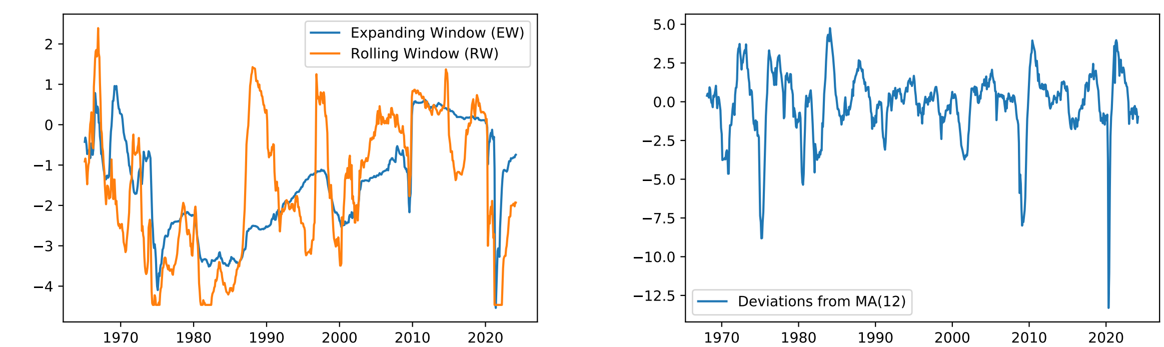
(a) Swaps Across Maturity (b) Comparison to Breakeven Inflation (TIPS)



The left figure presents a time series plot of the 1-year (blue), 5-year (orange), and 10-year (green) inflation swap rates. The right figure displays a time series plot of the 5-year zero-coupon inflation swap rate (blue) and the 5-year TIPS implied zero-coupon break-even inflation yield (orange). Yields are expressed as annual percentages.

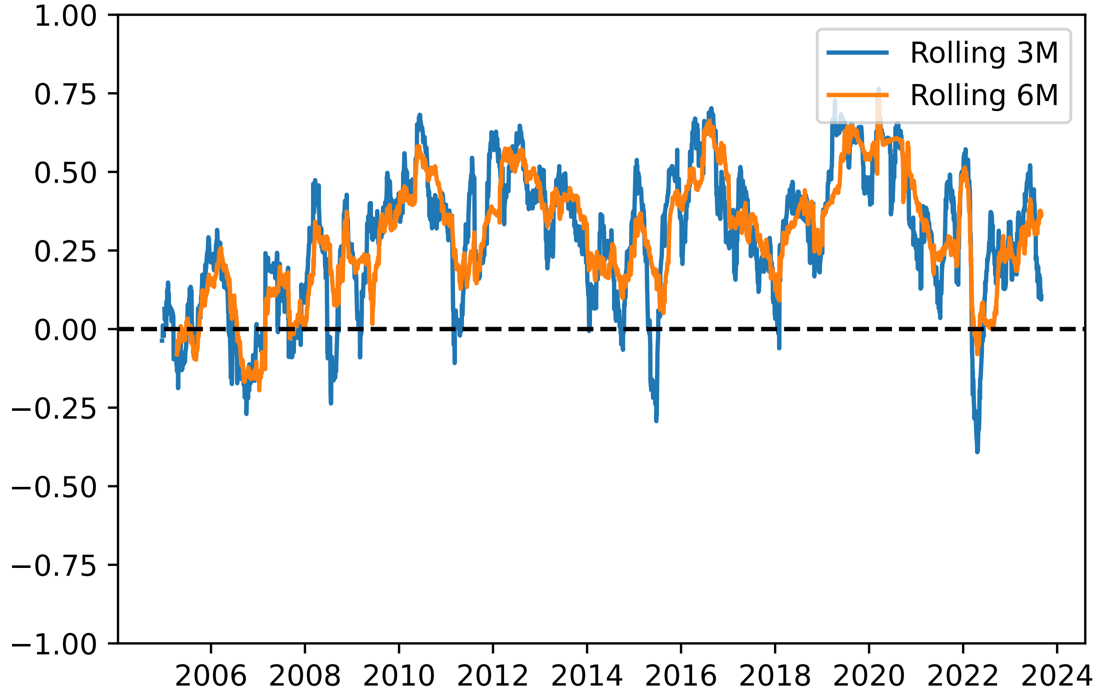
Figure A4: **Alternative Lower Frequency Macro Measures**

(a) Nominal-Real Covariance from Boons et al. (2020) (b) Capacity Utilization (Adjusted)



The left figure presents a monthly time series plot of the nominal real covariance computed on an expanding window through weighted least squares using exponential weights, identical to Boons et al. (2020) (blue) and a 60-month rolling window version of the same covariance (orange). The right figure displays an adjusted version of capacity utilization, constructed using deviations from a 12-month moving average.

Figure A5: **Inflation Swap and Market Return Correlation**



This figure presents a time series plot of the rolling 3-month (blue) and 6-month (orange) correlation between daily changes in 5-year inflation swap spreads and stock market returns.

Table A1: **Key Summary Statistics**

	Count	Mean	Std. Dev.	Min	Max
		(a) Aggregate Measures			
$\pi^{swap,1Y}$	730	1.903	1.168	-4.274	5.856
$\pi^{swap,5Y}$	730	2.222	0.533	-0.515	3.593
$\pi^{swap,10Y}$	734	2.423	0.379	0.992	3.190
$\Delta\pi^{swap,5Y}$	728	0.000	0.049	-0.285	0.191
$\rho\left(R_{bond}, R_{mkt}\right)^{3M}$	819	-0.293	0.280	-0.778	0.544
$\rho\left(R_{bond}, R_{mkt}\right)^{6M}$	819	-0.291	0.248	-0.733	0.433
$\rho\left(\Delta\pi^{swap}, R_{mkt}\right)^{3M}$	701	0.292	0.218	-0.348	0.746
$\rho\left(\Delta\pi^{swap}, R_{mkt}\right)^{6M}$	691	0.297	0.185	-0.167	0.704
		(b) Firm-Level Data			
$Spread$	418911	2.257	3.767	0.101	33.054
$\Delta Spread$ (b.p.)	418808	0.139	8.359	-52.475	65.279
$ExpLoss$	204936	0.639	1.529	0.029	14.191
$RiskPrem$	204757	1.206	1.922	-2.686	16.365
R_i (%)	207853	0.032	2.276	-9.615	9.253
$R_i - R_f$ (%)	207853	0.027	2.276	-9.619	9.250
		(c) Intraday Swaps			
$\Delta\pi^{idswap,5Y}$	622	0.116	3.364	-28.000	24.500
$\Delta\pi^{surp,5Y}$	622	0.052	1.208	-5.279	10.559
$\Delta\pi^{latent,5Y}$	622	0.097	2.703	-29.574	22.233

This table reports the aggregate measures and firm-level summary statistics for the variables used in the empirical analysis. Panel A reports aggregate measures on macroeconomic announcement days. Panel B reports summary statistics of firm-level CDS and equity returns on macroeconomic announcement days. Panel C reports summary statistics of intraday, 1-hour changes of 5Y inflation swaps surrounding macroeconomic announcements of interest. Subcomponents of the intraday changes are provided, based on the methodology from Gürkaynak et al. (2020). See main text for more details. CDS data come from Markit, and expected losses and risk premia are estimated using the conditional probability of default (EDF) and recovery rate estimates from Moody's Analytics and Markit, following Palazzo and Yamarthy (2022). Equity returns and excess returns come from CRSP. Intraday data are from Refinitiv TickHistory. All firm-level, daily data are winsorized at the 0.5% level.

Table A2: **Time-Varying Inflation Sensitivity of Equity Returns**

	(1)	(2)	(3)
$\Delta\pi^{swap,5Y}$	0.38***	0.35***	0.35***
	(3.91)	(3.82)	(3.92)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		0.05	
		(1.00)	
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		-0.22***	
		(-2.58)	
$\tilde{\rho}_{-1}^{bond-mkt,6M}$			0.07
			(1.59)
$\tilde{\rho}_{-1}^{bond-mkt,6M} \times \Delta\pi^{swap,5Y}$			-0.16**
			(-2.02)
$\left(R^i - R^f\right)_{-1}$	0.00	0.00	0.00
	(0.22)	(0.17)	(0.15)
$s_{i,-1}$	-0.00	-0.00	0.00
	(-0.10)	(-0.01)	(0.12)
Correlation Horizon	–	3M	6M
Firm FE	Y	Y	Y
Clustering		Firm-Time	
Obs	207,717	205,837	205,837
$Adj.R^2$	0.028	0.036	0.034

This table reports the sensitivities of equity markets to movements in inflation expectations. Column (1) reports the unconditional response of excess equity returns where inflation expectations are measured by inflation swaps. Column (2) reports results where inflation swap movements are interacted with the bond-stock correlation estimated using the 3-month rolling correlation, while column (3) uses the 6-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS spread and excess equity return the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level.* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A3: **Systematic Equity Returns and the Inflation-Growth Correlation**

	(1)	(2)	(3)	(4)	(5)
$\Delta\pi^{swap,5Y}$	0.35***	0.34***	0.03	0.38***	-0.00
	(3.82)	(3.57)	(1.55)	(3.80)	(-0.18)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$	0.05	0.04	0.01	0.06	-0.00
	(1.00)	(0.94)	(0.94)	(1.19)	(-0.25)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$	-0.22***	-0.22**	-0.02	-0.21**	-0.03**
	(-2.59)	(-2.47)	(-1.35)	(-2.25)	(-2.05)
$s_{i,-1}$	-0.00	0.01	-0.00	0.02	-0.01
	(-0.01)	(0.57)	(-0.16)	(0.84)	(-0.87)
$(R^i - R^f)_{-1}$	0.00	-0.01	0.02**	0.00	0.01
	(0.16)	(-0.36)	(2.19)	(0.01)	(1.10)
Risk Adjustment	–	Market Factor		Fama-French 3-Factor	
Dependent Variable	$R^i - R^f$	$(R^i - R^f)^{sys}$	$(R^i - R^f)^{idio}$	$(R^i - R^f)^{sys}$	$(R^i - R^f)^{idio}$
Firm FE	Y	Y	Y	Y	Y
Clustering	Firm-Time			Firm-Time	
Obs	205,721	205,721	205,721	205,721	205,721
$Adj.R^2$	0.036	0.083	0.002	0.074	0.002

This table reports time-varying sensitivities of equity returns, its systematic and idiosyncratic components to inflation expectation movements. Equity returns are decomposed into systematic and idiosyncratic components based on their risk factor loadings the day prior to macroeconomic announcements. We interact the inflation expectation movements with the bond-stock correlation estimated using the 3-month rolling correlation. Correlation measures are standardized such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. Column (1) reports the aggregate results, columns (2) - (3) focus on decomposing equity returns using the market factor, while columns (4) - (5) using Fama and French (1993) factors. In all regressions, we include either the CDS spread and equity return the day before the macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A4: **Inflation Sensitivity of Equity Returns Across Risk Groups**

	(1)	(2)	(3)	(4)
$\Delta\pi^{swap,5Y}$	0.35***	0.26***	0.34***	0.42***
	(3.82)	(3.47)	(3.70)	(3.60)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$	0.05	0.04	0.04	0.03
	(1.00)	(1.07)	(0.87)	(0.55)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$	-0.22***	-0.18**	-0.21**	-0.29***
	(-2.58)	(-2.47)	(-2.55)	(-2.79)
$s_{i,-1}$	-0.00	0.17	0.07	-0.00
	(-0.01)	(0.52)	(0.31)	(-0.10)
$\left(R^i - R^f\right)_{-1}$	0.00	-0.02	-0.02	0.03
	(0.17)	(-0.73)	(-0.74)	(1.44)
Which Risk Group	–	1	3	5
Firm FE	Y	Y	Y	Y
Clustering	Firm-Time			
Obs	205,837	41,453	41,166	40,862
$Adj.R^2$	0.036	0.044	0.043	0.029

This table reports time-varying sensitivities of equity returns to inflation expectation movements, across different risk groups. Firms are sorted into CDS risk quintiles based on 5-year CDS spreads the day prior to macroeconomic announcements. We interact the inflation expectation movements with the bond-stock correlation estimated using the 3-month rolling correlation. Correlation measures are standardized such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. The left most column reports the unconditional result, and the right three columns focus on risk groups 1, 3, and 5. In all regressions, we include the CDS spread and excess equity return on the day before macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A5: **Macro Surprises, Inflation Expectations, and Equity Returns**

	(1)	(2)	(3)	(4)
$\Delta\pi^{swap,5Y}$	0.38***	0.35***		
	(3.91)	(3.82)		
$\Delta\pi^{surp,5Y}$			0.01	0.07
			(0.11)	(1.09)
$\Delta\pi^{resid,5Y}$			0.39***	0.36***
			(3.91)	(3.94)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		-0.22***		
		(-2.58)		
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{surp,5Y}$				-0.18***
				(-3.84)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{resid,5Y}$				-0.14*
				(-1.77)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		0.05		0.06
		(1.00)		(1.28)
$s_{i,-1}$	-0.00	-0.00	-0.00	0.00
	(-0.10)	(-0.01)	(-0.12)	(0.02)
$\left(R^i - R^f\right)_{-1}$	0.00	0.00	0.00	0.00
	(0.22)	(0.17)	(0.22)	(0.11)
Firm FE	Y	Y	Y	Y
Clustering	Firm-Time			
Obs	207,717	205,837	207,717	205,837
$Adj.R^2$	0.028	0.036	0.030	0.043

This table reports time-varying sensitivities of equity markets to inflation expectation movements, focusing on the macro surprise and residual components of inflation expectations. Daily movements in inflation swaps are decomposed into surprise and residual components based on a regression procedure. We then interact each of these components with the bond-stock correlation estimated using the 3-month rolling correlation. Correlation measures are standardized such that the interaction coefficient indicates the additional sensitivity to changes in the surprise or residual component of inflation swaps when the correlation is one standard deviation higher. In all regressions, we include the CDS spread and equity return on the day before macroeconomic announcements, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A6: **Macroeconomic Announcements for Intraday Analysis**

Announcement	Time	Frequency	Observations	Unit	Std. Dev.
Core CPI	8:30	Monthly	184	% MoM	0.12
CPI	8:30	Monthly	184	% MoM	0.13
Nonfarm Payrolls	8:30	Monthly	196	Change	740.817k
GDP	8:30	Quaterly	54	% QoQ ann.	0.72
Core PPI	8:30	Monthly	188	% MoM	0.23
PPI	8:30	Monthly	188	% MoM	0.37

This table displays the selected macroeconomic announcements with their release times, frequencies, number of observations, units of measurement, and the conversion factor for a one standard deviation positive surprise to the original release unit. The data displays five major macroeconomic series examined throughout the paper, spanning from June 2007 to Oct 2023. "Frequency" denotes how often the data is released, while "Observations" refers to the total count of data points (surprises) for each macroeconomic series in the dataset. The term "Unit" indicates the measurement unit in which the data is reported.

Table A7: **Intraday Swap Prices and Macroeconomic Surprises**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\varepsilon^{corecpi}$	1.75***						0.91***
	(8.18)						(2.95)
ε^{cpi}		1.89***					1.28***
		(9.13)					(4.17)
$\varepsilon^{nonfarm}$			0.42**				0.45**
			(2.04)				(1.98)
ε^{gdp}				1.18			1.18***
				(1.47)			(2.71)
$\varepsilon^{coreppi}$					0.40**		0.13
					(2.00)		(0.45)
ε^{ppi}						0.54***	0.46
						(2.72)	(1.63)
Dependent Var.	Intraday $\Delta\pi^{swap,5y}$ (b.p.)						
Obs	184	184	196	54	188	188	622
$Adj.R^2$	0.265	0.310	0.016	0.022	0.016	0.033	0.120

This table reports the average effect of macroeconomic surprises on intraday inflation swap prices. Inflation swap data is collected daily from 8:15 AM to 9:15 AM ET, reflecting a 60-minute window. This table includes 622 announcements following October 2007. To ensure independence from monetary policy-related interest rate movements, days with FOMC announcements are excluded. Columns (1) - (6) report results of individual univariate regressions of intraday inflation swap movements onto macroeconomic surprises, while column (7) reports results of a multivariate regression including all macroeconomic surprises. Macroeconomic surprises are normalized by their respective standard deviations. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A8: **Latent Factor Estimation from Intraday Swaps**

	(1)	(2)	(3)	(4)	(5)	(6)
$\varepsilon^{corecpi}$	3.35***	2.79***	1.71***	0.90***	1.04***	0.65***
	(4.55)	(4.26)	(5.53)	(2.82)	(5.76)	(4.68)
ε^{cpi}	2.68***	2.41***	1.12***	1.30***	0.69***	0.79***
	(4.04)	(4.73)	(3.22)	(4.07)	(3.27)	(4.84)
$\varepsilon^{nonfarm}$	-0.11	0.01	0.06*	0.45***	0.38***	0.28***
	(-1.29)	(0.23)	(1.66)	(23.57)	(15.17)	(16.01)
ε^{gdp}	-0.19	-0.26	0.86	1.18***	-0.40	0.11
	(-0.23)	(-0.39)	(1.29)	(3.34)	(-1.08)	(0.42)
$\varepsilon^{coreppi}$	0.42	-0.71	0.73***	0.13	0.39***	-0.25
	(1.42)	(-0.98)	(2.78)	(1.19)	(2.61)	(-1.24)
ε^{ppi}	0.47**	0.41	0.48***	0.47***	0.44***	0.74***
	(2.34)	(1.42)	(2.92)	(3.56)	(3.27)	(3.28)
$\Delta\pi^{latent}$	2.56***	2.64***	3.46***	2.70***	2.33***	1.94***
	(4.09)	(6.32)	(21.15)	(29.57)	(17.21)	(16.23)
Dependent Variable	Intraday $\Delta\pi^{swap}$					
Horizon	1Y	2Y	3Y	5Y	7Y	10Y
Observations	622	622	622	622	622	622
R^2 without latent	0.235	0.208	0.119	0.120	0.091	0.096
R^2 with latent	0.410	0.434	0.769	0.771	0.665	0.709

This table reports the Kalman Filter estimates based on intraday data, as given in Equation 3. Inflation swap data is collected daily from 8:15 AM to 9:15 AM ET, reflecting a 60-minute window. This table includes 622 announcements or 6 relevant macroeconomic releases (corecpi, cpi, non-farm, gdp, coreppi and ppi) following October 2007. To ensure independence from monetary policy-related interest rate movements, days with FOMC announcements are excluded. Macroeconomic surprises are normalized by their respective standard deviations. The latent factor is estimated using changes in asset prices around macroeconomic releases similar to Gürkaynak et al. (2020). Each column reports results for a different maturity of intraday inflation swaps. The R^2 values are those of announcement day yields using (i) solely headline surprises vs. (ii) headline surprises and the latent factor. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A9: **PCA Decomposition of Inflation Expectations and Risk Premia**

(a) Correlations Across Principal Components

$\Delta\pi^{swap,5Y}$	1.0000				
$\Delta\pi^{be,5Y}$	0.6708	1.0000			
$\Delta yield^{real,5Y}$	-0.2326	-0.4186	1.0000		
Δ Risk Component	0.4378	0.1316	0.7654	1.0000	
Δ Expected Inflation Component	0.8308	0.9052	-0.6308	0.0000	1.0000

(b) Regressions PCA

	(1)	(2)	(3)	(4)	(5)	(6)
	Δ Inflation Swaps		Δ BE Inflation Rate		Δ Real Yield	
Expectation Component	3.937***	3.937***	4.063***	4.063***	-3.822***	-3.822***
	(90.42)	(146.42)	(129.07)	(135.72)	(-49.25)	(-300.52)
Risk Component		2.075***		0.590***		4.638***
		(77.17)		(19.73)		(364.66)
Obs	3,672	3,672	3,672	3,672	3,672	3,672
R^2	0.690	0.882	0.819	0.837	0.398	0.984

(c) Validation of PCA

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	KW TP	DKW Inf. RP	1-Year Ahead Inflation			5-Year Ahead Inflation		
Risk Component	1.640***	0.500***						
	(92.08)	(79.31)						
Expectation Component			1.083**			0.684*		
			(2.58)			(1.85)		
Inflation Swap				0.761			-0.031	
				(1.21)			(-0.38)	
Mich 1-year					0.499			
					(0.76)			
Mich 5-year								-2.343**
								(-2.26)
Obs	3,672	3,672	216	216	216	168	168	168
R^2	0.698	0.631	0.293	0.037	0.030	0.373	-0.006	0.347

This table reports the validations of the PCA procedure. The top panel reports correlations among the variables used in the PCA and the resulting components. The middle panel presents regression results of daily changes in inflation swaps, breakeven inflation, and real yields on the standardized principal components. The bottom panel reports validation regressions. Columns (1) and (2) report regressions of the term premium and D'Amico et al. (2018) inflation risk premium on the estimated inflation risk premium, while columns (3) - (8) report results of regressions of realized inflation at the 1 and 5 year horizons on the estimated expected component or inflation swaps and Michigan Surveys of Consumers expectations.

Table A10: **Time-Varying Inflation Sensitivities by Swap Maturity**

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\pi^{swap,1Y}$	-0.58***			0.04		
	(-4.27)			(0.34)		
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,1Y}$	0.43***			-0.04		
	(3.84)			(-0.35)		
$\Delta\pi^{swap,5Y}$		-0.81***			0.35***	
		(-5.27)			(3.82)	
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.61***			-0.22***	
		(5.05)			(-2.58)	
$\Delta\pi^{swap,10Y}$			-0.82***			0.39***
			(-6.21)			(4.96)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,10Y}$			0.58***			-0.25***
			(5.36)			(-3.49)
Dependent Variable		Δs_i			$R^i - R^f$	
Firm FE	Y	Y	Y	Y	Y	Y
Clustering		Firm-Time			Firm-Time	
Obs	411,822	410,129	412,898	206,956	205,837	207,638
$Adj.R^2$	0.018	0.024	0.025	0.002	0.036	0.049

This table reports the time-varying effects of inflation expectation movements, measured using different maturity of inflation swap rates, on movements in CDS and equity returns. Columns (1) - (3) focus on CDS movements, while columns (4) - (6) focus on equity returns. All columns report results where the inflation measures are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. Columns (1) and (4) report results using 1-year inflation swap rates, columns (2) and (5) report results using 5-year inflation swap rates, and columns (3) and (6) report results using 10-year inflation swap rates. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include either the CDS rate or the CDS rate and expected loss or the excess return the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A11: **Asset Price Response to Breakeven Inflation**

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\pi^{be,5Y}$	-0.99***	0.37***	-0.30***	-0.65***	-0.94***	0.35***
	(-6.68)	(4.80)	(-4.07)	(-5.05)	(-7.07)	(4.73)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$					0.02	0.04
					(0.29)	(0.83)
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{be,5Y}$					0.57***	-0.23***
					(4.99)	(-3.07)
$s_{i,-1}$	0.17***	-0.00	0.05	-0.00	0.17***	0.00
	(3.07)	(-0.14)	(1.42)	(-0.04)	(3.20)	(0.01)
$(R^i - R^f)_{-1}$		-0.00				-0.00
		(-0.01)				(-0.01)
$ExpLoss_{i,-1}$			-0.17***	0.55***		
			(-3.17)	(5.27)		
Dependent Variable	Δs_i	$R^i - R^f$	$\Delta ExpLoss_i$	$\Delta RiskPrem_i$	Δs_i	$R^i - R^f$
Firm FE	Y	Y	Y	Y	Y	Y
Clustering	Firm-Time		Firm-Time		Firm-Time	
Obs	440,133	223,199	210,332	210,330	432,551	221,319
$Adj.R^2$	0.020	0.028	0.009	0.012	0.025	0.038

This table reports the average and time-varying effects of inflation expectation movements, measured using 5-year TIPS breakeven inflation rates, on movements in CDS, expected losses, credit risk premia, and equity returns. Columns (1) and (5) focus on movements in CDS spreads, columns (2), and (6) on equity returns, and columns (3) and (4) on the expected loss component and credit risk premia, respectively. Columns (5) and (6) report results where the inflation measures are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include either the CDS rate or the CDS rate and expected loss or the excess return the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A12: **Time-Varying Inflation Sensitivities and CDS Liquidity**

	(1)	(2)	(3)
$\Delta\pi^{swap,5Y}$	-0.81***	-1.11***	-0.42***
	(-5.27)	(-5.47)	(-4.18)
$\tilde{\rho}_{-1}^{swap-mkt,3M}$	-0.03	-0.02	-0.03
	(-0.38)	(-0.20)	(-0.51)
$\tilde{\rho}_{-1}^{swap-mkt,3M} \times \Delta\pi^{swap,5Y}$	0.61***	0.78***	0.38***
	(5.05)	(5.12)	(4.45)
$s_{i,-1}$	0.18***	0.22***	0.14***
	(3.21)	(2.62)	(2.65)
Number of Dealers	–	High ($\geq$ 50%)	Low (< 50%)
Firm FE	Y	Y	Y
Clustering		Firm-Time	
Obs	410,129	234,586	175,517
$Adj.R^2$	0.024	0.037	0.020

This table reports the time-varying effects of inflation movements on CDS, controlling for CDS market liquidity. Column (1) reports the baseline effect using the full sample, while column (2) reports the time-varying effects where we focus on CDS contracts traded by a number of dealers larger than the sample median on an announcement day, and in column (3) we focus on CDS contracts traded by a number of dealers lower than the sample median. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS rate the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A13: **Time-Varying Inflation Sensitivities Including Non-Announcement Days**

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\pi^{swap,5Y}$	-0.68***	-0.64***	-0.64***	0.37***	0.35***	0.35***
	(-9.76)	(-10.02)	(-10.21)	(10.27)	(10.42)	(10.68)
$\tilde{\rho}_{-1}^{bond-mkt,3M}$		-0.05*			0.01	
		(-1.85)			(0.85)	
$\tilde{\rho}_{-1}^{bond-mkt,3M} \times \Delta\pi^{swap,5Y}$		0.20***			-0.07***	
		(4.18)			(-2.76)	
$\tilde{\rho}_{-1}^{bond-mkt,6M}$			-0.06**			0.02
			(-2.13)			(1.20)
$\tilde{\rho}_{-1}^{bond-mkt,6M} \times \Delta\pi^{swap,5Y}$			0.19***			-0.06**
			(3.87)			(-2.35)
$s_{i,-1}$	0.06***	0.06**	0.06**	0.00	0.00	0.00
	(2.65)	(2.38)	(2.36)	(0.38)	(0.52)	(0.59)
$\left(R^i - R^f\right)_{-1}$				-0.02*	-0.02*	-0.02*
				(-1.66)	(-1.65)	(-1.66)
Dependent Variable		Δs_i			$R^i - R^f$	
Correlation Horizon	–	3M	6M	–	3M	6M
Firm FE	Y	Y	Y	Y	Y	Y
Clustering		Firm-Time			Firm-Time	
Obs	2,822,262	2,700,173	2,700,173	1,373,117	1,353,636	1,353,636
$Adj.R^2$	0.011	0.012	0.012	0.027	0.028	0.028

This table reports the sensitivities of credit and equity markets to movements in inflation expectations including non-announcement days. Columns (1) and (4) report the unconditional response, for CDS spreads and equities, respectively. Columns (2) and (5) report results where the inflation swap movements are interacted with the bond-stock correlation estimated using the 3-month rolling correlation, while columns (3) and (6) use the 6-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS spread or CDS spread and equity returns the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A14: **Time-Varying Inflation Sensitivities over a Long Sample**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta\pi^{InflComp}$	-0.134***		-0.277***		-0.138***		-0.246***	
	(-3.746)		(-7.438)		(-3.647)		(-6.918)	
$\Delta\pi^{ExpInfl}$		-0.182***		-0.392***		-0.190***		-0.353***
		(-3.601)		(-7.443)		(-3.552)		(-6.983)
$\Delta\pi^{InflComp} \times \tilde{\rho}_{-1}^{bond-mkt,3M}$	0.383***							
	(4.198)							
$\Delta\pi^{ExpInfl} \times \tilde{\rho}_{-1}^{bond-mkt,3M}$		0.577***						
		(4.462)						
$\Delta\pi^{InflComp} \times \tilde{\rho}_{-1}^{bond-mkt,6M}$					0.322***			
					(3.222)			
$\Delta\pi^{ExpInfl} \times \tilde{\rho}_{-1}^{bond-mkt,6M}$						0.481***		
						(3.415)		
$\mathbb{1}_{\rho_{-1}^{bond-mkt,3M}>0} \times \Delta\pi^{InflComp}$			0.230***					
			(3.385)					
$\mathbb{1}_{\rho_{-1}^{bond-mkt,3M}>0} \times \Delta\pi^{ExpInfl}$				0.325***				
				(3.379)				
$\mathbb{1}_{\rho_{-1}^{bond-mkt,6M}>0} \times \Delta\pi^{InflComp}$							0.171**	
							(2.288)	
$\mathbb{1}_{\rho_{-1}^{bond-mkt,6M}>0} \times \Delta\pi^{ExpInfl}$								0.253**
								(2.408)
Correlation Horizon	3 Months				6 Months			
Obs	1,338	1,338	1,338	1,338	1,338	1,338	1,338	1,338
Adj.R^2	0.043	0.044	0.038	0.038	0.037	0.038	0.034	0.034

This table reports the time-varying effects of inflation compensation and expectations on ICE BofA US Corporate Index Option-Adjusted Spread from 1996 to 2023. All inflation measures come from D'Amico et al. (2018), where inflation compensation is defined as the sum of inflation expectations and inflation risk premia. Columns (1) - (4) report results where the inflation shocks are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. Columns (5) - (8) report results where the inflation shocks are interacted with the bond-stock correlation estimated using 6-month rolling correlation. Columns (3) - (4) report results where the inflation expectation movements are interacted with a dummy variable, that indicates whether the 3-month bond-stock correlation (non-standardized) is positive. Columns (7) - (8) report results where the inflation expectation movements are interacted with a dummy variable, that indicates whether the 6-month bond-stock correlation (non-standardized) is positive. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation measures when the correlation is one standard deviation higher. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A15: **Time-Varying Inflation Sensitivities over a Long Sample - Equity**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta\pi^{InflComp}$	0.063**		0.341***		0.051		0.275***	
	(2.073)		(5.541)		(1.599)		(4.974)	
$\Delta\pi^{ExpInfl}$		0.069**		0.352***		0.056*		0.284***
		(2.151)		(5.360)		(1.653)		(4.817)
$\Delta\pi^{InflComp} \times \tilde{\rho}_{-1}^{bond-mkt,3M}$	-0.281***							
	(-9.645)							
$\Delta\pi^{ExpInfl} \times \tilde{\rho}_{-1}^{bond-mkt,3M}$		-0.288***						
		(-9.577)						
$\Delta\pi^{InflComp} \times \tilde{\rho}_{-1}^{bond-mkt,6M}$					-0.246***			
					(-7.798)			
$\Delta\pi^{ExpInfl} \times \tilde{\rho}_{-1}^{bond-mkt,6M}$						-0.252***		
						(-7.683)		
$\mathbb{1}_{\rho_{-1}^{bond-mkt,3M}>0} \times \Delta\pi^{InflComp}$			-0.536***					
			(-8.023)					
$\mathbb{1}_{\rho_{-1}^{bond-mkt,3M}>0} \times \Delta\pi^{ExpInfl}$				-0.551***				
				(-7.835)				
$\mathbb{1}_{\rho_{-1}^{bond-mkt,6M}>0} \times \Delta\pi^{InflComp}$							-0.464***	
							(-7.523)	
$\mathbb{1}_{\rho_{-1}^{bond-mkt,6M}>0} \times \Delta\pi^{ExpInfl}$								-0.476***
								(-7.369)
Correlation Horizon	3 Months				6 Months			
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
Clustering	Firm-Time				Firm-Time			
Obs	7,259,306	7,259,306	7,259,306	7,259,306	7,259,306	7,259,306	7,259,306	7,259,306
$Adj.R^2$	0.014	0.014	0.012	0.012	0.012	0.012	0.010	0.011

This table reports the time-varying effects of inflation compensation and expectations on equity returns from 1983 to 2023. All inflation measures come from D'Amico et al. (2018), where inflation compensation is defined as the sum of inflation expectations and inflation risk premia. Columns (1) - (4) report results where the inflation shocks are interacted with the bond-stock correlation estimated using the 3-month rolling correlation. Columns (5) - (8) report results where the inflation shocks are interacted with the bond-stock correlation estimated using 6-month rolling correlation. Columns (3) - (4) report results where the inflation expectation movements are interacted with a dummy variable, that indicates whether the 3-month bond-stock correlation (non-standardized) is positive. Columns (7) - (8) report results where the inflation expectation movements are interacted with a dummy variable, that indicates whether the 6-month bond-stock correlation (non-standardized) is positive. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation measures when the correlation is one standard deviation higher. In all regressions, we include the equity returns the day before the macroeconomic announcement and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A16: **Time-Varying Inflation Sensitivities and the Inflation Swap-Market Correlation**

	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\pi^{swap,5Y}$	-0.90***	-1.02***	-1.03***	0.38***	0.46***	0.47***
	(-5.19)	(-6.33)	(-6.04)	(3.91)	(4.97)	(4.92)
$\tilde{\rho}_{-1}^{swap-mkt,3M}$		-0.20**			0.02	
		(-2.58)			(0.55)	
$\tilde{\rho}_{-1}^{swap-mkt,6M}$			-0.20**			0.04
			(-2.55)			(0.84)
$\tilde{\rho}_{-1}^{swap-mkt,3M} \times \Delta\pi^{swap,5Y}$		-0.68***			0.38***	
		(-5.55)			(5.73)	
$\tilde{\rho}_{-1}^{swap-mkt,6M} \times \Delta\pi^{swap,5Y}$			-0.56***			0.34***
			(-4.56)			(4.91)
$s_{i,-1}$	0.18***	0.19***	0.19***	-0.00	0.00	-0.00
	(3.12)	(3.21)	(3.25)	(-0.10)	(0.07)	(-0.04)
$\left(R^i - R^f\right)_{-1}$				0.00	-0.01	-0.00
				(0.22)	(-0.33)	(-0.08)
Dependent Variable		Δs_i			$R^i - R^f$	
Correlation Horizon	–	3M	6M	–	3M	6M
Firm FE	Y	Y	Y	Y	Y	Y
Clustering		Firm-Time			Firm-Time	
Obs	418,777	405,195	400,641	207,717	202,603	199,661
$Adj.R^2$	0.019	0.026	0.024	0.028	0.056	0.049

This table reports the time-varying effects of inflation expectation movements on credit and equity markets, using a correlation measure based on daily movements of swap rates and aggregate equity returns. Columns (1) and (4) report the baseline, unconditional results. Columns (2) and (5) report results where the inflation expectation movements are interacted with the 3-month swap-market correlation, while columns (3) and (6) use the 6-month rolling correlation. We standardize the correlation measures such that the interaction coefficient indicates the additional sensitivity to changes in inflation swap when the correlation is one standard deviation higher. In all regressions, we include the CDS rate or CDS rate and equity returns the day before the macroeconomic announcement, and firm fixed effects. Standard errors are clustered at the time and firm level. * Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

Table A17: **Baseline Model Calibration**

(a) Model Parameters

	Value	Notes
γ	20	Bansal and Shaliastovich (2013)
ψ	2.5	Target risk-free rate
δ	0.998	Bansal and Shaliastovich (2013)
μ_c	0.00474	Target consumption growth mean
μ_π	0.009	Bansal and Shaliastovich (2013)
Π_{cc}	0.95	Bansal and Yaron (2004)
$\Pi_{\pi\pi}$	0.988	Bansal and Shaliastovich (2013)
σ_{xc}	0.0000583	Target expected growth vol
$\sigma_{x\pi}$	0.000986	Target expected inflation vol
$\sigma_{xc\pi}(s_1)$	0.0008	"Good Inflation" regime
$\sigma_{xc\pi}(s_2)$	-0.0004	"Bad Inflation" regime
p_{11}	0.9	–
p_{22}	0.9	–
σ_c	0.00359	Target consumption growth vol
σ_π	0.00557	Target inflation vol
$\beta_{\lambda 0}$	0.00505	Target 2% annual default rate
$\beta_{\lambda xc}$	-0.5	Countercyclical default rates
R	0.4	Average recovery rate from Markit

(b) Model-Implied Values

	Value	Notes
$E\left[pc_t\right]$	7.607	Log price-consumption ratio
$E\left[r_{ct}\right]$	2.011	Real return on consumption
$E\left[r^{\$}_{ct}\right]$	5.538	Nominal return on consumption
$E\left[r^{\$}_{ft}\right]$	4.629	Nominal risk-free rate
$E\left[r_{ct} - r_{ft}\right]$	0.908	Risk premium
$E\left[r^{5Y,\$}_{ft}\right]$	3.466	Nominal return on 5Y risk-free bond
$E\left[s^{5Y}_t\right]$	1.337	5Y CDS spread
$\sigma\left[\Delta s^{5Y}_t\right]$ (b.p.)	5.371	Volatility of spread changes
$\rho(r^{\$}_{ct}, r^{5Y,\$}_{ft})$	-0.148	Bond-stock correlation
$\rho(r^{\$}_{ct}, r^{5Y,\$}_{ft})$ – Regime 1	-0.451	–
$\rho(r^{\$}_{ct}, r^{5Y,\$}_{ft})$ – Regime 2	0.284	–
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$	0.231	Excess return regression coefficient
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$ – Regime 1	0.933	
$\beta(r_{ct} - r_{ft} \sim \Delta x_{\pi t})$ – Regime 2	-0.475	
$\beta(\Delta s^{5Y}_t \sim \Delta x_{\pi t})$ (b.p.)	-1.603	Spread change regression coefficient
$\beta(\Delta s^{5Y}_t \sim \Delta x_{\pi t})$ – Regime 1	-6.265	
$\beta(\Delta s^{5Y}_t \sim \Delta x_{\pi t})$ – Regime 2	3.073	

This table presents parameters used to calibrate the model and the simulated model implied values. The top panel shows the baseline parameters. Some parameters come from the literature, while parameters related to consumption growth and inflation are calibrated using 1968Q4 to 2019Q4 data. The bottom panel displays the results of the model simulation, where we simulate 40,000 quarters, including a burn-in period.