Inflation Risk and Yield Spread Changes[*]

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ABSTRACT

Inflation risk explains about 50% of the systematic variation of yield spread changes beyond standard structural factors. Unexpectedly high inflation decreases real debt value, reducing real liabilities and firms' default probabilities. The effect is heterogeneous and non-linear, increasing in firm leverage and credit rating and decreasing in maturity and ex-ante inflation rate, consistently with a default risk channel. Accounting for non-linearities, inflation risk reduces the explanatory power of the residual systematic component of yield spread changes by 18.2 percentage points.

JEL classification: G10; G12; G20

Keywords: Corporate Bonds, Inflation Risk, Yield Spread Changes, Inflation Linked-derivatives.

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I. Introduction

Changes in corporate yield spreads are empirically challenging to explain. Collin-Dufresne et al. (2001) (hereafter CDGM) shows that conventional structural variables play a role, yet a substantial amount of unexplained common variation remains. U.S. corporate debt is overwhelmingly nominal, thus, an increase in the inflation rate erodes the real value of debt and decreases firms' default risk. Therefore, a natural conjecture is that inflation risk constitutes a substantial piece of the puzzling unexplained common variation.

In this paper, we empirically investigate the ability of inflation risk to explain the large excess common variation in yield spread changes. We find that inflation risk accounts for about 50% of the systematic variation of yield spread changes beyond standard structural variables. The effect is heterogeneous and non-linear, increasing in firm leverage and credit rating and decreasing in maturity and in ex-ante inflation rate. Further, we find that, accounting for non-linearities, inflation risk reduces the explanatory power of the residual systematic component of yield spread changes by 18.2 percentage points.

Following CDGM and, more recently, Friewald and Nagler (2019) and He et al. (2022) (hereafter FN and HKS respectively), we run individual-bond time series regressions of yield spread changes on CDGM structural factors, using our sample of 5905 bonds from 2005 to 2021. Our transaction data reveal similar properties compared to CDGM, FN and HKS, that is, we find the explanatory power of monthly yield spread changes to be low, with a mean adjusted R^2 value of 33.3% and, using principal component analysis (PCA), we find that 80.2% of the residual variation can be explained by the first principal component, indicating a large systematic component not captured by structural credit variables.

After confirming in our sample the puzzles of low explanatory power and large systematic component of yield spread changes, we connect inflation risk to the common variation of yield spread changes. Following the recent literature on inflation and deflation risks (Haubrich et al. (2012), Fleckenstein et al. (2016) and Fleckenstein et al. (2017)) we exploit a simple market-based approach to measure inflation risk. That is, we employ zero-coupon inflation swaps, forward contracts whereby the inflation buyer pays a predetermined fixed nominal rate and in return receives from the seller an inflation-linked payment. Zero-coupon inflation swaps are one of the most liquid of all over-the-counter inflation-linked derivative products and, together with nominal treasuries, provide an alternative measure of real yields (Fleming et al. (2013)). Unlike Treasury Inflation-Protected Securities (TIPS), zero-coupon inflation swaps do not display embedded deflation options or liquidity issues, making their rates a more reliable measure of inflation. We obtain closing bid and ask quotes of inflation swaps for each trading day from Bloomberg for annual maturities from one to ten years, as well as for 12-, 15-, 20-, and 30-year maturities. We adjust the swap rates for seasonal patterns and we match the cash-flow structure of each bond by performing a spline interpolation between provided maturities whenever necessary. We then use the matched swap rates to compute our measures of inflation risk.

To account for the multiple facets of inflation risk, we define several proxies. Theories of Chen (2010), Kang and Pflueger (2015), Gomes et al. (2016) and Bhamra et al. (2022), relate credit spread to inflation risk through the nominal cash flow channel. These theories predict that an increase in inflation decreases the real value of the debt, and since firms do not adjust their leverage immediately, their default risk, and thus their yield spreads decrease. We use the changes in cash-flow matched swap rate as a measure of changes in inflation rate. Furthermore, to capture potential nonlinear effects due to convexity, we also include the squared change in cash flow matched swap rate. In yield spread regressions we find that both measures are significant, exhibit the predicted sign, and explain 18.4% and 24.1% of the total variation of residuals of yield spread changes, respectively.

David (2008) relates inflation uncertainty to the uncertainty about the current fundamental values. As in Merton (1974), defaultable bond price is modeled as a risk-free bond price minus the price of the put option on the nominal asset value of the firm; therefore, an increase in inflation uncertainty increases the value of this put option, increasing the bond credit risk. In a similar fashion, in Kang and Pflueger (2015), an increase in inflation uncertainty raises defaults and credit losses through its effect on firms' default thresholds. Furthermore, Fischer (2016) suggests that inflation uncertainty may affect bonds' value through the delay of investment due to uncertainty in prices. To capture the different aspects of inflation uncertainty, we compute a short-term and a long-term measure, such as the 21 and 120 trading days standard deviations of the cash-flow-matched swap rate. Both proxies should be positively correlated with yield spreads as all channels suggest an increase in overall riskiness in the firm's debt due to the higher ex ante probability of default. Overall, short-term and long-term inflation volatilities explain 21.1% and 17.9% of the total variation of residuals of yield spread changes, respectively.

Lastly, the slope of the inflation curve embodies expected future inflation rates, which could have an effect on yield spreads. A steepening of the curve indicates higher future spot rates, i.e., higher future inflation rate, which could be beneficial depending on the current health of the economy. A raise in future expected inflation rate is favorable for firms if the economy is moving out from a low-inflation—deflationary state, while it could harmful if it occurs with ex-ante high inflation rate. In fact, an increase in the expected inflation rate when the ex-ante level is high could indicate a future recession, which would lead to an increase in the yield spreads (Gilchrist and Zakrajšek (2012)). On the other hand, inflation risk has a marginally decreasing effect on yield spreads, despite which, the former effect will dominate regardless (Bhamra et al. (2022)). To capture this effect, we use the slope of the swap curve, defined as the difference between the 10-year and the 2-year swap rate.

We then proceed to investigate all the proxies jointly. All proxies are statistically significant with t-statistics ranging from 8.48 to 36.93. Yield spreads narrow with increases in the inflation rate, and widen with increases in convexity, slope of the swap curve, short-term and long-term volatilities. Overall, inflation risk increases the mean and median adjusted R^2 values by 12.8 and 12.4 percentage points, respectively; it accounts for 49.9% of the unexplained variation of yield spread changes, and it also reduces the explanatory power of the latent factor by 11.5 percentage points.

Motivated by the large effect found, we investigate whether inflation has a heterogeneous impact on yield spread changes. Theories of Kang and Pflueger (2015) and Bhamra et al. (2022) justify the effect of inflation on bonds mainly through default risk; therefore, an increase in inflation will be more beneficial to riskier firms than to relatively safer firms. Similarly, inflation risk could have a different impact depending on the maturity of the bonds. While long-term bonds accrue effects of changes in inflation rates for a longer time span, long-term inflation rates are sticky and do vary considerably less than its short-term counterparts. To test the heterogeneous effects of inflation risk, we run time-series regressions of yield spread changes on our proxies and compute average coefficients and their statistical significance differentiating among cohorts of leverage, rating and time to maturity. We find that inflation risk's effect is increasing in firm's leverage and credit rating, consistent with a default risk explanation, and is decreasing in time to maturity, consistent with inflation risk being anchored at longer horizons.

Next, we analyse the non-linearities of inflation risk. Bhamra et al. (2022) argues that increases in inflation have a more pronounced effect on default risk when it is currently high, which is when inflation is low. The marginally decreasing effect of inflation risk might bias our estimates and mislead our interpretation. We solve this issue by interacting the inflation proxies with a dummy for high inflation periods and estimating inflation risk effects of low/medium and high inflation periods separately. We find that inflation risk display a relevant non-linear dimension, which, in some proxies, is large enough to suppress the baseline effect. Accounting for the non-linearities further increases the explanatory power of inflation risk on yield spread changes, reaching an increase of 15.8 and 16 percentage points in the mean and median adjusted R^2 values, respectively, and further decreases the fraction of variance explained by the common component, achieving a 18.2 percentage points decrease.

In additional tests, we show that the significant decrease in explanatory power of the first principal component due to inflation risk does not depend on how we aggregate residuals. In fact, results are similar when residuals are aggregated by cohorts based on time to maturity and leverage, ratings, dollar trading volume of the past month, the beta of the bonds' stock market, or VIX betas. We show that our results are not driven by bonds of firms with sticky cash-flows. Then, we replicate the baseline results using TIPS' break-even rates in place of the inflation swaps rates, showing that our results are not driven by the specific inflation-linked instrument we use. Lastly, we demonstrate the robustness of our results by controlling for alternative variables related to yield spread changes. We control for changes in cash-to-market value, inflation volatility risk of Ceballos (2021), changes in unemployment and real consumption and income, OTC market proxies of Friewald and Nagler (2019) and intermediary distress of He et al. (2022).

This article primarily contributes to the empirical literature linking inflation to asset prices. In the equity market, the relevance of inflation risk has been extensively documented (Fama (1981), Chen et al. (1986), Weber (2014), Eraker et al. (2016), Fleckenstein et al. (2017), Boons et al. (2020)). In the corporate bond market, the recent literature has attempted to provide evidence on the relevance of inflation risk. Kang and Pflueger (2015) documents that inflation volatility and inflation cyclicality have a significant impact on aggregate credit spreads for a panel of developed economies. Bhamra et al. (2011), and Gomes et al. (2016) study the effect of long-term nominal debt as a transmission mechanism for inflation via a sticky leverage channel. Illeditscha (2018) states that the component of inflation risk correlated with real assets and risky cash flows is priced in corporate bonds. Ceballos (2021) finds a negative inflation volatility risk premium obtained from the difference between high inflation and low inflation beta portfolios, and Augustin et al. (2021) explores the relation between price rigidity and credit risks in the cross section. We differ and contribute to this literature by studying yield spread changes and linking the residual variation with inflation risk.

We also contribute to the credit risk modelling literature. The unexplained common variation of yield spread changes, first documented in Collin-Dufresne et al. (2001) and most recently studied by Friewald and Nagler (2019) and He et al. (2022), is a canonical puzzle in the context of structural models like Merton (1974), Leland (1994) and Leland and Toft (1996). As well related is the "credit spread puzzle" of Huang and Huang (2012). Theories of Chen (2010), Kang and Pflueger (2015), Gomes et al. (2016) and Bhamra et al. (2022), relate credit spread to inflation risk through the nominal cash flow channel. While, the work of David (2008) and Fischer (2016) relate inflation uncertainty to uncertainty about current fundamental values and to delay of investment due to uncertainty in prices. We contribute to this literature by providing empirical support for theories that incorporate inflation risk into structural models.

Further, our paper contributes to the literature of inflation linked securities and their application. Pflueger and Viceira (2011) documented a relative high correlation between TIPS's and nominal bonds over short investment horizons. Fleming et al. (2013) studies the trading activity and liquidity in the inflation swap market. Haubrich et al. (2012) develops a model of nominal and real bond yield curves and estimates it with inflation swaps. Fleckenstein et al. (2014) shows that Treasury bonds are consistently overpriced relative to inflation-swapped TIPS, while D'Amico et al. (2018) investigates the poorer liquidity of TIPS relative to nominal Treasury securities. Christensen et al. (2016) and Fleckenstein et al. (2017) study the nature of deflation risk using inflation swaps and options. We add to this literature by developing inflation risk proxies from inflation swap rates.

The remainder of the paper is organized as follows. Section II describes the data and Section III presents the baseline analysis, replicating CDGM. Section IV introduces our measures of inflation risk. Section V presents the main results using our inflation risk proxies to explain yield spread changes. Section VI concludes the paper.

II. Data

We rely on several data sources to analyze the impact of inflation risk on yield spread changes. The sample of corporate bond transactions comes from the enhanced Trade Reporting and Compliance Engine (TRACE) maintained by the Financial Industry Regulatory Authority (FINRA). We follow the cleaning steps from Dick-Nielsen (2014), thus cleaning same-day corrections and cancellations, removing reversals, and double counting of agency trades. Then, we follow Edwards et al. (2007) and apply a median filter and a reversal filter to eliminate further potential data errors. Whereas the median filter identifies potential outliers in reported prices within a certain time period, the reversal filter identifies unusual price movements[1]. The sample period is 2005 to 2021. We merge corporate bond pricing data with the Mergent Fixed Income Securities database (FISD) to obtain bond characteristics, such as offering amount, offering date, maturity, coupon rate, bond rating, bond option features, and issuer information, and with CRSP/Compustat data using the linking table provided by the Wharton Research Data Services (WRDS). Following the literature related to corporate bonds, we restrict the sample to corporate debentures and exclude bonds that have variable coupons, are convertible, putable, asset backed, exchangeable, privately placed, perpetual, preferred securities, secured lease obligations, unrated, or quoted in a foreign currency. We also remove bonds issued by financial firms (Standard Industrial Classification, or SIC, codes 6000–6999) or utility firms (SIC codes 4900–4999) and bonds with issue sizes under $10 million or a time-to-maturity of more than 30 years or less than one month.

We follow CDGM and obtain market-and firm-specific variables that, according to structural models, determine yield spread changes. In particular, we obtain market variables such as the Standard & Poor's (S&P) 500 index from the Center for Research in Security Prices (CRSP), the volatility index (VIX) from the Chicago Board Options Exchange, and

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[1] The median filter eliminates any transaction where the price deviates by more than 10% from the daily median or from a nine-trading-day median centered at the trading day. The reversal filter eliminates any transaction with an absolute price change deviating from the lead, lag, and average lead/lag price change by at least 10%.

the 10-year Treasury constant maturity rate from daily off-the-run Treasury yield curves constructed by Gürkaynak et al. (2007). As a systematic proxy for the probability or magnitude of a downward jump in firm value, we construct a measure based on at- and out-of-the money put options and at- and in-the-money call options with maturities of less than one year, traded on the SPX index. We obtain option-implied volatilities from OptionsDX. For the exact procedure for estimating the jump component, we refer to CDGM. We use market leverage as a proxy for a firm's creditworthiness. Following FN we define market leverage as book debt over the sum of book debt and the market value of equity, where book debt is given by the sum of Compustat items Long-Term Debt - Total (DLTT) and Debt in Current Liabilities - Total (DLC). To account for varying time lags between a firm's fiscal year-end and the information becoming publicly available, we apply a conservative lag of six months before we update a firm's debt-related information. The market value of equity is the number of common shares outstanding times the share price, both obtained from the CRSP. The inflation data comes from two sources. We obtained daily bid and ask quotes for the inflation swap from Bloomberg for annual maturities of 1 to 10 years, as well as for 12-, 15-, 20- and 30-year maturities, from July 2004[2]. Zero-coupon TIPS yields and break-even rates are obtained from Gürkaynak et al. (2008), who derive them from TIPS coupon bond yields, for annual maturities from 2 to 19 years.

The main variable in the empirical analysis is the yield spread. Using TRACE intra-day data, we first eliminate transactions with when-issued, lock-in, special trades or primary trades flag. Then we calculate the daily clean price as the volume-weighted average of intraday prices to minimize the effect of bid-ask spreads in prices, following Bessembinder et al. (2009). We then consider the closest observation to the last trading day of the month, within a five trading days window, as the month-end observation. We compute the end-of-month corporate bond yield from the volume weighted price and define the yield spread as the

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[2] Bloomberg does not retain quotes for inflation swaps before July 23, 2004, even though their trading started earlier. The 2- to 10-year swap maturities started trading in April 2003; the 12-, 15- and 20-year inflation swap rates start in November 2003; and the 30-year inflation swap rates start in March 2004.

difference between the bond yield and the yield of a risk-free bond with the same cash flow structure as the corporate bond. We use the U.S. Treasury yield curve estimates obtained from the Federal Reserve Board as risk-free benchmark. Next, we compute the monthly changes and returns of all the variables. To avoid asynchronicity issues, we match the dates of any variable available at daily frequency (e.g., VIX) to the dates in which we measure the end-of month bond prices. Following CDGM, we only consider bonds for which we have at least 25 observations of monthly yield spread changes.

[Insert Table I here.]

Table I reports the summary statistics of the sample of corporate bonds. The sample consists of 435,602 observations of monthly yield spread changes of 5905 bonds issued by 912 firms. The average yield spread is 2.41%, with a standard deviation of 3.24%. The average offering size is 748.47 millions of dollars and the average time-to-maturity is 8.99 years. Around 24% of the observations are high-yield bonds. We determine a bond rating as the average of ratings provided by Standard & Poor (S&P), Moody's and Fitch Ratings when at least two ratings are available or as the rating provided by one of the three rating agencies when only one rating is available.

III. CDGM Determinants of Yield Spread Changes

We first replicate Table 2 in CDGM and show that the large commonality in yield spread residuals persists in our sample of the U.S. corporate bond market from 2005–2021. Following CDGM, we use the same firm-specific and macroeconomic determinants of yield spread changes, motivated by structural models ‘a la Black and Scholes (1973) and Merton (1974). Specifically, we use changes in firm leverage ($\Delta Lev_{i,t}$), changes in 10-year Treasury interest rate (ΔRF_t), squared changes in the 10-year Treasury interest rate (ΔRF_t^2), changes in the slope of the yield curve ($\Delta Slope_t$), measured as the difference between 10-year and 2-year Treasury interest rates, changes in the VIX index (ΔVIX_t), S&P 500 return (RM_t), and changes in a jump factor ($\Delta Jump_t$) based on S&P 500 index options. We define the vector of CDGM of bond i at time t as:

$$\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]. \tag{1}$$

We estimate the following regression model for each bond i with yield spread changes $\Delta YS_{i,t}$:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \varepsilon_{i,t} \tag{2}$$

Following CDGM and FN, we assign each bond to a leverage group based on the firm's average leverage ratio over the bond's lifetime. The groups are defined as below 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and above 55%. We present the average coefficients and their statistical significance for each cohort in Table II. The associated t-statistics are calculated from the cross-sectional variation over the coefficient estimates within a cohort. Thus, for each cohort we divide the average coefficient by the standard deviation of the coefficient estimates and scale by the square root of the number of bonds in the cohort. To facilitate the presentation of the results, in Table II column (7) we also present the average coefficients and their statistical significance in a regression where we use all 6676 bonds. The signs of the coefficients are economically meaningful and comparable with CDGM, that is, yield spreads increase with firm leverage, the slope of the term structure, volatility and jump risk, and decrease with the risk-free rate and the market return. The mean and median adjusted R^2 are low, as pointed out by CDGM and FN, with an overall mean value of 33.3%.

[Insert Table II here.]

Then we analyze the commonality in the yield spread residuals. Following the empirical procedure of CDGM, we assign each bond into one of 18 cohorts based on time-to-maturity (under five years, five to eight years, and more than eight years) and leverage (as defined above). For each of the 18 cohorts we compute the average of the regression residuals $\boldsymbol{\varepsilon}_{g,t}$

across bonds i in cohort g in month t, and then conduct a PCA on these residual series to capture the properties of the remaining variation. We present the results in Table III.

[Insert Table III here.]

Importantly, there is a strong systematic factor structure of the regression residuals. The total unexplained variance is 140 basis points, of which 80.2% is explained by the first principal component, PC1, while the second component, PC2, accounts for only 4.2%. These results are very much in line with the findings of CDGM and HKS, which report an explanatory power of the first PC of 75% and 80% respectively, while being significantly larger than FN, with only 48.4%. Following HKS, the last column of Table III reports the common variation of residuals of each cohort g, σ_{ε_g} $(=\sum_t \varepsilon_{g,t} - \hat{\varepsilon}_{g,t})$ as a fraction of the total variation of the 18 cohorts $(\sigma_{\varepsilon_g}/\sum_g^{15} \sigma_{\varepsilon_g})$. As in HKS, the above 55% leverage group accounts for the majority of variation, in fact, summing across all maturities, it accounts for 42.1% of the overall variation.

Importantly, we confirm in our sample that the explanatory power of monthly yield spread changes is very low, with a mean adjusted R^2 value of 33.3% and, using principal component analysis (PCA), that the residuals exhibit high commonality, i.e. the fraction of total unexplained variance of the regression residuals that can be accounted for by the first principal component is 80.2%, similar to previous studies.

IV. Inflation Proxies

In this section, we first motivate the use of zero-coupon inflation swaps through a simple nominal rate decomposition and then establish our inflation risk proxies. Subsequently, we use these measures to investigate the ability of inflation risk to explain the variation of yield spread changes, by first analyzing the effect of each proxy separately within the framework of CDGM, and then considering their joint impact. Next, we investigate the heterogeneous effect by looking across leverage, ratings and time to maturity cohorts, then we consider their ability to explain changes in credit spreads when their non-linearities are taken into account. Lastly, we show that our results are robust to different sorts of bonds regression residuals.

A. Simple Nominal Rates Decomposition

The nominal yield on a zero-coupon bond of maturity m, $y^n_{t,m}$ can be decomposed into a real yield, $y^r_{t,m}$ (the yield on a zero-coupon bond perfectly indexed against inflation), plus inflation compensation, $\theta_{t,m}$. Inflation compensation reflects both expected inflation $E_t[\pi_{t,m}]$, and an inflation risk premium $\phi_{t,m}$ (ignoring Jensen's inequality terms):

$$\begin{aligned} y^n_{t,m} &= y^r_{t,m} + \theta_{t,m} \\ &= y^r_{t,m} + E_t[\pi_{t,m}] + \phi_{t,m} \end{aligned} \tag{3}$$

Using inflation protected bonds or inflation-linked derivatives, it is possible to empirically achieve the decomposition of nominal yields, $y^n_{t,m}$, into real yield and inflation compensation ($y^r_{t,m}$, $E_t[\pi_{t,m}]$, and $\phi_{t,m}$).

In 1997, the U.S. Treasury started issuing Treasury Inflation-Protected Securities (TIPS), fixed coupon bonds whose principal amount is adjusted daily based on the third preceding calendar month CPI for All Urban Consumers. Beginning with the first TIPS auction, market participants began making markets in inflation swaps as a way of hedging inflation risk. Zero-coupon inflation swaps are among the most liquid of all over-the-counter market inflation derivative products. These swaps are forward contracts whereby the inflation buyer pays a predetermined fixed nominal rate and in return receives from the seller an inflation-linked payment. They are quoted with maturities ranging from one to 30 years and together with nominal Treasuries, they provide an alternative measure of real yields.

In this study, we will employ inflation swap rates as a more reliable reflection of inflation compensation for multiple reasons. First, TIPS inflation adjustment is bounded below at its issuance value. Thus, TIPS offer an embedded put option that protects investors against deflation on the bond's principal payment (Grishchenko et al. (2016) and Christensen et al. (2016)). Since this option has a non-negative value, its presence increases a TIPS' price, and hence makes TIPS' yield lower than a yield of a bond perfectly indexed against inflation. Zero-coupon inflation swap contracts do not contain this option. Therefore, all else equal, break-even inflation rate (Treasury rate minus TIPS rate) based on a TIPS principal strip should be higher than the equivalent-maturity inflation swap rate. Second, studies by Elsasser and Sack (2004), Fleckenstein et al. (2014), D'Amico et al. (2018) and Andreasen et al. (2021) reveal that the TIPS break-even inflation rates consistently fell below survey measures of inflation expectations and that Treasury bonds are almost always overvalued relative to an inflation swapped TIPS. This mispricing narrows as additional capital flows into the markets and as market liquidity increases. Thus, TIPS' yields contain a large liquidity premium, due to the fact that TIPS, like other bonds, tend to go into buy-and-hold investors' portfolios as time passes, which makes break-even inflation rates differ even further from inflation compensation rates.

[Insert Figure 1 here.]

Figure 1 shows the time series of 2-year zero-coupon treasury yield, break-even, and inflation swaps in the top panel, while in the bottom panel the difference between the 2-year zero-coupon inflation swap rate and the 2-year TIPS' implied zero-coupon break-even inflation yield. Generally, the pattern is consistent with previous studies; the inflation swap rate minus TIPS' implied break-even rate exhibits time variation, with a positive average and a peak during the financial crisis. This evidence is consistent with Campbell et al. (2009) and Haubrich et al. (2012), which attribute the spike in TIPS yields following Lehman Brothers' bankruptcy to Lehman's use of substantial amounts of TIPS to collateralize its repo borrowings and derivative positions, and with Fleckenstein et al. (2014) which finds that the price difference narrows when the US auctions nominal Treasuries or TIPS, and it widens when dealers have difficulty obtaining Treasury securities, such as during a period of increased repo failures.

B. *Inflation Proxies*

Zero-coupon inflation swaps are available from 1 to 30 year maturities and are quoted daily. We obtain closing bid and ask quotes of inflation swap rates for each trading day from Bloomberg for annual maturities from one to ten years, as well as for 12-, 15-, 20-, and 30-year maturities[3]. Since both inflation swaps and TIPS are indexed to the seasonally unadjusted CPI, we adjust their rates following Fleckenstein et al. (2014). We first estimate seasonal weightings for the CPI-U for each month of the year by regressing the CPI-U index values for the January 1980 to December 2021 period on monthly indicator variables. The estimated weights are normalized to ensure that there is no seasonal effect for full-year swaps and then used to adjust the interpolated inflation swap curve (seasonal adjustments are not used for maturities less than one year)[4]. We detail the full procedure in the Appendix. We then match the cash-flow structure of each bond and obtain cash-flow structure matched swap rates by performing a spline interpolation between provided maturities whenever necessary. We use the cash-flow matched rates to compute the following proxies:

1. *Swap Rate Level.* Theories of Chen (2010), Kang and Pflueger (2015), Gomes et al. (2016) and Bhamra et al. (2022), relate yield spread to inflation risk through the nominal cash flow channel. These theories predict that an increase in inflation decreases the real value of the debt, and since firms do not adjust their leverage immediately, their default risk and thus their yield spreads decrease. We use the changes in cash-flow matched swap rate ($\Delta Swap_{i,t}$) as a measure of changes in inflation rate. Furthermore, to capture potential nonlinear effects due to convexity, we also include the squared change in cash flow matched swap rates ($\Delta Swap^2_{i,t}$). According to theory, we expect

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[3] We disregard other maturities because are less liquid and their quotes appear to be stale.

[4] We begin our seasonal adjustment with the shortest available maturity, hence 1-year for the zero-coupon inflation swaps rates and 2-years for the TIPS break-even rates.

changes in inflation compensation to be negatively associated with changes in credit spread.

2. *Short-term Volatility of Swap Rate.* David (2008) relate inflation uncertainty to uncertainty about current fundamental values. As in Merton (1974), defaultable bond price is modeled as a risk-free bond price minus the price of the put option on the nominal asset value of the firm; thus, an increase in inflation uncertainty increases the value of this put option, increasing the bond credit risk. In a similar fashion, in the model of Kang and Pflueger (2015), an increase in inflation uncertainty raises defaults and credit losses through its effect on firms' default thresholds. To capture these effects, we define the short-term volatility of swap rate as the 21 trading days standard deviation of the cash flow matched swap ($\Delta\sigma^{21}i,t$). According to both theory models, this proxy should be positively correlated with yield spreads.

3. *long-term Volatility of Swap Rate.* The model of Fischer (2016) suggests that inflation uncertainty can affect the value of bonds by delay in investment due to uncertainty in prices. To capture the long-term uncertainty of inflation, we define the long-term volatility of swap rates as the 120 trading days standard deviation of the cash flow matched swaps ($\Delta\sigma^{120}i,t$). Following Fischer (2016), long-term inflation uncertainty should be positively correlated with yield spreads.

4. *Slope of Swap Curve.* The slope of the inflation curve embodies expected future inflation rates which could have an effect on yield spreads. A steepening of the curve indicates higher future spot swap rates, i.e., higher future inflation rate, which could be beneficial depending on the current health of the economy. For example, a raise in future expected inflation rate is beneficial for a firm if the economy is moving out from a low inflation/deflationary state, while it could harm the firm if it occurs when inflation rates are already high. The rationale being that an increase in the expected inflation rate when its level is already high could indicate a future recession, which will increase credit spreads. Furthermore, expected inflation decreases default marginally

less when inflation is already high, not fully compensating the increase from expected recession (Bhamra et al. (2022). To capture these effects, we use the slope of the swap curve, defined as the difference between the 10-year and 2-year swap rates ($\Delta Slope^S_t$). We do not have a clear prediction for the effect of swap curve on the yield spreads.

In Table IV, we report the unconditional correlations between the changes in yield spreads and our proxies, as well as among the proxies.

[Insert Table IV here.]

When comparing our measures, we find that, generally, the absolute correlation coefficients are relatively low, when compared to similar measures obtained from nominal treasury rates. In fact, the highest pairwise correlation of 58.6% is between $\Delta Slope^S_S$ and $\Delta Swap_{i,t}$, while the similar correlation between ΔRF_t and $\Delta Slope_t$ is 66.3%. These low correlations prompt the use of all our proxies in the further empirical analysis and suggest that each measure reflects a slightly different aspect of the corresponding inflation risk. Table IV also reports the standard deviations of all our variables to ease the interpretation of the economic impact of our proxies in the subsequent regression analyses.

V. Inflation Risk and Yield Spread Changes

A. *Baseline*

We examine the impact of inflation risk on yield spread changes by augmenting the CDGM baseline specification by our measures:

$$\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap^2_{i,t},\ \Delta\sigma^{21}i,t,\ \Delta\sigma^{120}i,t,\ \Delta Slope^S_t]. \tag{4}$$

We run the following time-series regression for each bond i:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\Delta\beta}^T_i \boldsymbol{S}_{i,t} + \boldsymbol{\Delta\theta}^T_i \boldsymbol{I}_{i,t} + \nu_{i,t}, \tag{5}$$

and report the average coefficients across bonds, t-statistics, and the mean and median R^2 in Panel A of Table V. Column (1) reiterates the results of the CDGM baseline model. We test each of the proxy separately in columns (2) to (6), and then jointly in column (7).

[Insert Table V here.]

All proxies are statistically significant when both added separately and jointly, with t-statistics ranging from 8.48 to 36.93. Yield spreads narrow with increases in swap rates and widen with increases in convexity, slope of the swap curve, and short-term and long-term volatilities. Each of the proxies, on average, increases the mean adjusted R^2 value between 3.2 and 6.4 percentage points compared to the baseline specification. Employing all variables together increases the mean and median adjusted R2 values by 12.8 and 12.4 percentage points, respectively. Note that this is a sizable fraction, given that we explain changes in yield spreads and not their levels.

Next, to evaluate the overall explanatory power of the inflation risk proxies on yield spread changes, we first repeat the exercise in Table III, by computing, for each of the 18 cohorts, the average of the regression residuals $\nu_{g,t}$ across bonds i in cohort g in month t when a new variable is added, and then conduct a PCA on these residual series to capture the properties of the remaining variation. Table V Panel B reports the results. We find that inflation proxies decrease the proportion of unexplained variance associated with the common component, PC1, by 4.9 percentage points on average, while the overall decrease is 11.5 percentage points, that is, from 80.2% in the CDGM benchmark to a value of 68.7%. Furthermore, following HKS, we compute the fraction of the total variation of residuals that is accounted for by each new variable. In particular, we compute the total unexplained variation of the yield spread residuals after adding each proxy, σ_ν $(= \sum_t \nu_{g,t} - \hat{\nu_{g,t}})$, and then we compute the fraction of variation explained as

$$FVE = 1 - \frac{\sum_g^{18} \sigma_{\nu_g}}{\sum_g^{15} \sigma_{\varepsilon_g}}, \tag{6}$$

where g are again the 18 cohorts sorted by leverage and maturity. Each of the inflation risk proxy reduces the unexplained variance by 25.8 basis points, or by 18.4 percentage points, while overall inflation risk accounts for 49.9% of the total variation of the residuals of changes in yield spread. This explanatory power is large, even when directly compared to previous studies. At a quarterly frequency, HKS shows that dealer inventory and an intermediary distress factor explain 43% of the total variation of residuals, while FN shows that OTC market frictions explain 45% of the total variation of residuals.

Although we have shown that the coefficients of our proxies are statistically significant and that their explanatory power is large, their economic importance also merits discussion. We rely on the full model and analyze the implied yield spread change for a one standard deviation change in a particular measure. For example, $\Delta Swap_{i,t}$ has a pricing impact of close to seven basis points, while $\Delta Swap^2_{i,t}$ and $\Delta Slope^S_t$ are around four basis points. The short-term volatility ($\Delta\sigma^{21}i,t$) has the second largest price impact of around six basis points, while the long-term volatility ($\Delta\sigma^{120}i,t$) has the smallest impact of around three basis points. The economic pricing impacts obtained are quite substantial, considering that, in the full model, the pricing impact of one standard deviation change in 10-years treasury rate is also around seven basis point, which is close to the pricing impact of the change in swap rate. Furthermore, the price impact is also substantial when compared to mean and median yield spreads. In fact, a one standard deviation change in the swap rate decreases the mean yield spread of 3.43% and the median of 5.4%.

In sum, our baseline analysis shows that (i) inflation risk has a significant effect on yield spreads, (ii) the five proxies, having low correlation, together account for more than half of the unexplained variation of yield spread changes, and (iii) a substantial proportion of the latent factor is related to time-varying inflation risk.

B. Heterogeneity

Theories of Kang and Pflueger (2015) and Bhamra et al. (2022) justify the effect of inflation on bonds mainly through default risk, hence an increase in inflation will be more beneficial to riskier firms than relatively safer firms. Similarly, inflation risk could have a different impact depending on the maturity of the bonds. While long-term bonds accrue effects of changes in inflation rates for a longer time span, long-term inflation rates are sticky and do vary considerably less than its short-term counterparts. Furthermore, the increase in explanatory power also depends on the inflation risk proxies underlying effect, since in Table 2 we show that mean and median adjusted R^2 vary between groups. To test the inflation risk heterogeneous effects, we run time-series regressions following Eq. 5, and we compute average coefficients and their statistical significance differentiating among cohorts. Table VI presents the results by average firm leverage during the life of the bond, where the groups are defined as below 15%, 15%– 25%, 25%–35%, 35%–45%, 45%–55%, and above 55%.

[Insert Table VI here.]

This table is equivalent to Table II with the addition of the inflation risk proxies, and thus each column should be compared with the relative column in Table II. All coefficient magnitudes are monotonically increasing in leverages, consistent with the notion of inflation risk effect being dependent on the ex-ante default risk of the firm. Bonds of more leveraged firms will be affected to a greater extent by inflation risk. Intuitively, an increase in inflation will lead to a larger decrease in the real debt-to-equity ratio in highly levered firms, since they have a higher nominal debt-to-equity ratio. This greater decrease in real leverage causes the yield spread of highly levered firms to shrink further. The mean adjusted R^2 increase on average by 11.9 percentage points while the median R^2 by 11.2 percentage points, with closely related increases ranging from 10.8 to 13.8 percentage points. The explanatory power increase is similar to the full sample R^2 increase.

Next, we study another aspect of bond's heterogeneity and default risk: bonds' credit ratings. We run time-series regression following Eq. 5, and we compute average coefficients and their statistical significance by rating groups, where the groups are defined as having a rating of AAA-AA, A, BBB, BB and B-C. To guarantee consistency across results, we consider bonds with at least 25 monthly observations within rating cohorts. Table VII presents the results.

[Insert Table VII here.]

For comparison, the odd-numbered columns report the result of the CDGM baseline model following Eq. 2, while even numbered columns reports the model including inflation risk proxies following Eq. 5. The rating cohorts are not evenly distributed, with the BBB cohort having the largest number of bonds with 2377 bonds and the AAA-AA cohorts exhibiting the lowest number of bonds with only 538. All proxies coefficient magnitudes monotonically increase with bond ratings, except $\Delta Swap^2_{i,t}$, which coefficient peak in the BBB cohort and then decrease, even though its statistical significance is low. The coefficients increasing in bond's rating are again consistent with inflation risk being dependent on the ex-ante default risk. Low-rated bonds have higher default risk and larger yield spreads; an increase in inflation will lead to a larger decrease in default risk for low-rated than for high-rated bonds. In terms of explanatory power, the mean adjusted R^2 increase on average by 10.2 percentage points while the median R^2 by 9.9 percentage points, with a wide range of values going from 6.1 to 14.6 percentage points. The greatest increase is in the most populated cohort, the BBB cohorts, with a 15.7% percentage points increase in mean adjusted R^2 and 14.6% in mean adjusted R^2.

Lastly, we explore the maturity dimension of bonds' heterogeneity. Hence, as before, we run time-series regression following Eq. 5, and we compute average coefficients and their statistical significance by maturity groups, where the groups are defined as less than five years, five to twelve years, and over twelve years, and we consider bonds with at least 25 monthly observations within maturity cohorts. Table VIII presents the results.

[Insert Table VIII here.]

Columns (1), (3) and (5) reiterate the results from Eq. 2 in the sub-samples, while columns (2), (4) and (6) reports the results including inflation risk proxies. In all columns, the magnitude of coefficients is similar, with long-term bonds exhibiting marginally larger coefficients. This could be miss-interpreted as inflation risk being more pronounced in long-term bonds. However, when we consider the average and standard deviation of the inflation proxies, the result is reversed. In fact, since we use cash-flow matched proxies, the short-term proxies have, on average, 37 and 3.3 times larger mean and standard deviations, respectively, than their long-term equivalent. Hence, a one standard deviation increase in the swap rate leads to 12.39 basis points decrease in short-term bonds yield spread, while only 5.8 basis points in yield spread of long-term bonds. If we then consider the different average yield spreads across cohorts, the same change in swap rate leads to an 6.23% decrease in average yield spread in short-term bonds, while on 2.34% in long-term bonds. The largest increase in R^2 appears also in the short-term cohort, with increments of 13.3 and 11.9 percentage points in mean and median adjusted R^2, respectively. The other cohorts show smaller, yet significant, increases, while overall, across the three cohorts, the increments average 8.9 and 7.8 percentage point in mean and median adjusted R^2.

Overall, we find evidence of heterogeneous effects of inflation risk on yield spreads. We show that the effect is increasing in firm leverage and credit rating, consistent with a default risk explanation, and is decreasing in time to maturity, consistent with inflation risk being anchored at longer horizon. We further show that the effect and the increase in explanatory power is not concentrated in a specific group of bonds.

C. Non-Linearities

We now investigate the non-linear effect of inflation on yield spread changes. For this purpose, we run the following time-series regression for each bond i:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} \times H_{t-1} + \nu_{i,t}, \tag{7}$$

where $\boldsymbol{S}_{i,t}$ is the vector of the structural model variables defined in Section III, the vector $\boldsymbol{\Delta I}_{i,t}$ refers to the proxies for inflation risk introduced in Section IV, and $\boldsymbol{H}_{t-1}$ is a dummy variable taking the value of 1 when the 1-year inflation swap rate in month $t-1$ is above the 80 percentile[5]. We define months in which $\boldsymbol{H}_{t-1}$ is 1 as ”high” inflation months. We impose more filtering on the sample to correctly estimate the interaction coefficients. Namely, we require each bond to have at least 6 monthly observations during ”high” inflation months and likewise outside ”high” inflation periods. Since most of the ”high” inflation states are at the beginning and toward the end of the sample, the additional filtering reduces the number of observations by 38.6%. While the decrease is significant, the filtered sample still covers the full period and has similar distributions of ratings, yield spreads, and offering amounts as the full sample. A modest difference can be found in terms of time to maturity, as the filtered sample mechanically includes fewer short-term bonds, the average time to maturity is 1.8 years longer. This is likely to bias the explanatory power downward, as in Table VIII the most incremental effect is in terms of R^2 comes from the short-term cohort.

[Insert Table IX here.]

Table IX reports the results. In column (1) we present the CDGM baseline following Eq. 2. When compared with the full sample, the restricted sample shows an higher mean and median adjusted R^2, respectively of 37.6% and 39.6%, and a larger total unexplained variance of 162 bps, of which 71.3% is explained by the first principal component; the PC1 explanatory power in this case is slightly lower than in the full sample. In columns (2) to (6),

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[5] Using 2-, 5-, or 10-years inflation swaps rates results in qualitatively and quantitatively similar results.

we estimate each proxy individually with their interaction term; by estimating two slopes, we allow inflation risk to play a different role depending on whether or not inflation was high. All proxy baseline coefficients are significant, with the same sign as the baseline case, while their magnitudes are larger. The interaction terms, except $\Delta Swap^2_{i,t}$ one, are of opposite sign with respect to the baseline proxies, consistent with inflation risk playing different roles on yield spreads depending on the ex-ante level. Most noticeable, the $\Delta Swap_{i,t}$ effect is canceled out when ex-ante inflation is high, following the notion that a further increase in inflation harms the firm when its default risk is already low, that is, when inflation is high. Inflation volatilities and slope changes also report similar results. In columns (6) and (7) we estimate the join effect both without and with the interaction terms. In the case without non-linearities, inflation risks explains 43.5% of the total variation of residuals of yield spread changes, while decreasing the PC1 by 15.9 percentage points. When non-linearities are taken into account, the fraction of variance explained is 48.9%, while the decrease in PC1 is 18.2 percentage points, respectively, an 12.4% and 14.4% increase due to the non-linearities.

Altogether, inflation risk exhibits a relevant non-linear dimension, which, in some proxies, is large enough to suppress the baseline effect. Accounting for non-linearities further increases the explanatory power of inflation risk on yield spread changes and further decreases the fraction of variance explained by the common component.

D. Additional Evidence and Robustness

In this section, we demonstrate that the results are not driven by the specific way to aggregate residuals and that our measures of inflation risk convey additional information for yield spread changes across different industry cash-flow stickiness groups. Further, we show that the results are invariant to the use of TIPS' break-even rates to create inflation proxies and to alternative variables related to yield spread changes.

A concern with our previous analysis is that the significant decrease in explanatory power of the first principal component due to inflation risk depends on the specific way to aggregate residuals. In this section, we show that our results are independent of the specific residual cohorts. In fact, we follow HKS and aggregate the time-series residuals according to five different cohorts. First, we use our baseline definition of time-to-maturity, leverage, and ratings cohorts following the same definition as in Tables V and VII. Second, we create quintile sorts based on past month dollar trading volume, bonds' stock market beta, and VIX betas. The past month trading volume is the sum of all trades in each bond in the previous month, while the betas of the stock market and VIX are the regression betas on the S&P 500 and the VIX in Eq. 2. Then, for each bond we run time-series regression following the CDGM model in Eq. 2. We assign the regression residuals to different cohorts based on time to maturity and either leverage, volume, stock market beta, or VIX beta. Depending on the groups, these allocations will result in 18 or 15 cohorts. We then compute an average residual for each cohort and extract the principal components of the covariance matrix of the residuals. Subsequently, we repeat the same procedure while using Eq. 5 including all inflation risk proxies. Table X reports the results.

[Insert Table X here.]

For each pair of grouping variables, we report the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the fraction of variance explained as defined in Eq. 6. The first row reports the results of the baseline model, while the second row reports the results of the model that include inflation risk proxies. Panel A reports results excluding non-linearities; The first two rows are identical to Table V. The mean reduction in the explanatory power of PC1 in all cohorts is 11.7 percentage points, in line with the 11.5 percentage points in Table V. The time-to-maturity and volume sort shows the least increase in explanatory power, with only 8.3 percentage points, and the time-to-maturity and VIX beta sort exhibits the largest increase, with 14.3 percentage points. The average fraction of variance explained (FVE) is 51.6%, with a very narrow range. In Panel B, we report results including non-linearities, where the sample is the same as in Table IX. Again, the first two rows reiterate the results of Table IX. The mean reduction in the explanatory power of the first principal component is 16.1 percentage points, slightly lower than the 18.2 percentage points in Table IX. The difference between the average of the five groups and the baseline scenario is due to the time to maturity and volume sort being low, with only 10.7 percentage points decrease, while all the other sorts exhibit at least a 15 percentage points decrease. As before, the time to maturity and VIX beta sort display the largest decreases, with 19.4 percentage points. The average fraction of variance explained (FVE) is 50.4%, with a moderately wider range than in Panel A.

Next, we show that our results are not driven by bonds of firms with sticky cash-flows. We define the cash-flow stickiness as the average absolute variation of the industry's producer price index (PPI), during the life of each bond. The PPI data come from the U.S. Bureau of Labor Statistics. In Table IA.I we presents time-series regressions of yield spread changes onto inflation proxies where average coefficients and their statistical significance are calculated within the PPI groups of the industry. We consider bonds with at least 25 monthly observations within PPI cohorts. We assign each bond to a cohort based on the average absolute PPI of the bond and report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. Across all the cohorts, inflation risk has a significant effect and the increase in explanatory power is not concentrated in a specific group.

Then, we show that our results are not driven by the specific inflation-linked instrument we use. We replicate the baseline results of Table V using TIPS break-even rates instead of the inflation swap rates. Table IA.II presents time-series regressions of yield spread changes onto inflation proxies based on off-the-run TIPS rates. For each of the inflation proxies, the sign, coefficient, and magnitude and the overall increase in explanatory power and decrease in the explanatory power of the first PC are similar to the baseline case.

Last, we demonstrate the robustness of our results by controlling for alternative variables related to yield spread changes. In Table IA.III, we run time-series regressions of yield spread changes onto inflation proxies, controlling for changes in cash to market value, inflation volatility risk (IVR) from Ceballos (2021) and changes in unemployment and real consumption and income. While, in Table IA.IV, we control for OTC market proxies from Friewald and Nagler (2019) and intermediary distress from He et al. (2022). Overall, the additional tests confirm that inflation risk is a major determinant of the dynamics of yield spread changes.

VI. Conclusion

We empirically study whether inflation risk drives the large unexplained common variation in yield spread changes. Inflation risk accounts for approximately 50% of the systematic variation of yield spread changes beyond standard structural factors. Unexpectedly high inflation decreases real debt value, reducing real liabilities and firms' default probabilities. The effect is heterogeneous and non-linear, increasing in firm leverage and credit rating and decreasing in maturity and ex-ante inflation rate, consistent with the default risk channel. When non-linearities are allowed, inflation risk reduces the explanatory power of the systematic component by 19.1 percentage points.

Appendix

A. Seasonal Adjustment of Swap and TIPS rates

Both TIPS and zero coupon inflation swaps are indexed to the non-seasonally adjusted U.S. CPI index, hence seasonal patterns in inflation must be taken into account when matching with corporate bond cash flows for swap maturities that include fractional years (e.g., 7.5 years). We adjust swap rates and TIPS yield following the procedure of Fleckenstein et al. (2014). Specifically, we first fit a standard cubic spline through the quoted maturities of both swaps and TIPS using a grid size of one month. We then estimate seasonal components in inflation from the monthly non-seasonally adjusted U.S. CPI index (CPI-U NSA) series between January 1980 and December 2021 by estimating an OLS regression of monthly log changes in the CPI index on month dummies. Thus, we obtain an estimate of the seasonal effect in each month. We normalize these seasonal factors so that their product is unity, thus, by construction, there will be no seasonal adjustment for full-year maturities. Next, we construct monthly forward rates from the interpolated rates, and multiply the forward rates by the corresponding adjustment factor. Lastly, we obtain the seasonally adjusted rates by converting the forward rates into spot rates. We do not interpolate or adjust maturity shorter than the quoted ones, i.e. one year for the swaps and two years for the TIPS, but instead use the shortest quoted maturity rate. The last step is as suggested by Fleckenstein et al. (2014), since the interpolated rates would then be sensitive to short-term inflation assumptions.

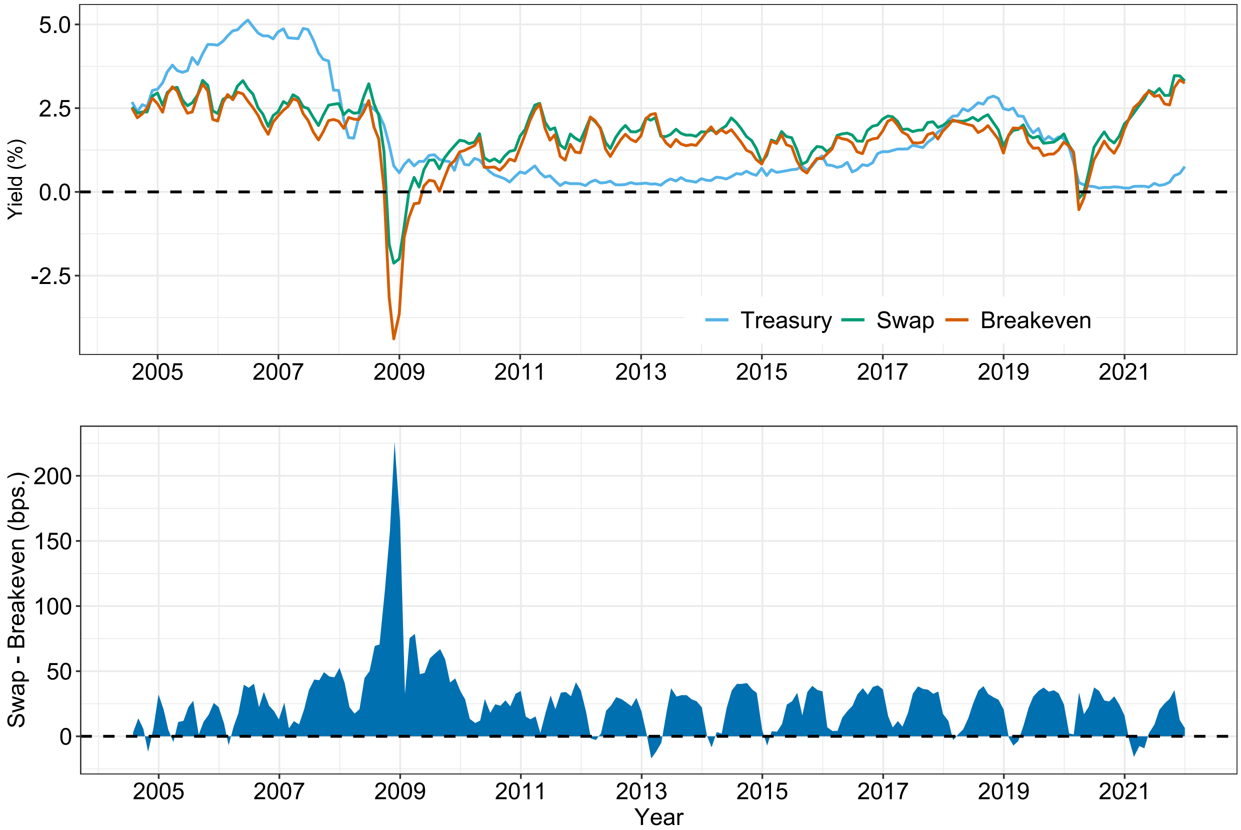

Figure 1. Time Series of Treasury, TIPS and Swaps rates
The top panel shows the time series of 2 year zero coupon treasury yield, break-even and inflation swaps. The bottom panel represent difference between the 2 year zero coupon inflation swap rate and the 2 year TIPS implied zero coupon break-even inflation yield. Yields are expressed as annual percentages, and the difference is in annual basis points.

Table I
Summary Statistics

This table reports summary statistics of the data. We report the number of observations, the mean, the standard deviation, and 5%, 25%, 50%, 75% and 95% quantiles of bond characteristics and the yield spread. The bond characteristics comprise the offering amount, the coupon rate, the bond age, the time to maturity, the duration, and the credit rating, where we assign integer numbers to the credit ratings (i.e., AAA=1, AA+=2,...) The bond's rating is determined as the average of ratings provided by Standard & Poor (S&P), Moody's and Fitch when more than one are available or as the rating provided by one of the three rating agencies when only one rating is available. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	Obs.	Mean	Std.	5%	25%	50%	75%	95%
Offering amount (mil.)	435,602	748.47	662.94	200	350	500	1,000	2,000
Coupon (%)	435,602	5.35	1.96	2.25	3.88	5.38	6.75	8.50
Age	435,602	5.49	5.31	0.52	2.04	3.97	7.02	17.37
Time to Maturity	435,602	8.99	8.09	0.96	3.37	6.14	9.99	27.30
Duration	435,602	6.50	4.31	1.00	3.32	5.45	8.24	15.59
Rating	435,602	8.75	3.48	4	6	9	10	15
Yield spread (%)	435,602	2.41	3.24	0.31	0.83	1.51	2.81	6.75
$\Delta YS_{i,t}$	435,602	0.89	85.58	-66.02	-15.38	-1.12	13.26	68.95

Table II
Determinants of Yield Spread Changes in Collin-Dufresne et al. (2001)

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \varepsilon_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section II, with $\Delta Lev_{i,t}$ as the change in firm leverage, ΔRF_t the change in 10-year Treasury interest rate, ΔRF_t^2 the squared change in the 10-year Treasury interest rate, $\Delta Slope_t$ the change in the slope of the term structure, ΔVIX_t the change in ΔVIX_t index, RM_t the S&P 500 return, and $\Delta Jump_t$ the change in a jump factor based on S&P 500 index options. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	<15%	15%–25%	25%–35%	35%–45%	45%–55%	>55%	All
Intercept	0.012	0.015	0.032	0.036	0.071	0.123	0.040
	(4.699)	(7.643)	(9.421)	(7.995)	(9.160)	(11.356)	(19.514)
$\Delta Lev_{i,t}$	0.012	0.006	0.014	0.021	0.032	0.058	0.020
	(2.762)	(5.729)	(10.807)	(12.715)	(9.596)	(14.545)	(17.671)
ΔRF_t	-0.312	-0.429	-0.562	-0.815	-0.853	-1.263	-0.632
	(-25.015)	(-33.892)	(-25.775)	(-21.840)	(-12.944)	(-16.822)	(-43.549)
ΔRF_t^2	0.039	0.093	-0.059	0.026	-0.213	-0.128	-0.009
	(0.831)	(4.055)	(-1.162)	(0.416)	(-2.182)	(-0.986)	(-0.345)
$\Delta Slope_t$	0.297	0.416	0.513	0.788	0.770	0.971	0.565
	(15.314)	(22.326)	(15.532)	(14.081)	(8.159)	(9.120)	(27.638)
ΔVIX_t	0.004	0.005	0.008	0.009	0.010	0.006	0.006
	(2.042)	(6.326)	(4.962)	(4.925)	(3.040)	(1.690)	(7.927)
RM_t	-0.016	-0.023	-0.039	-0.056	-0.082	-0.123	-0.048
	(-7.518)	(-25.030)	(-21.667)	(-21.554)	(-18.145)	(-24.660)	(-42.538)
$\Delta Jump_t$	0.002	0.004	0.006	0.012	0.012	0.022	0.008
	(2.250)	(8.194)	(6.396)	(8.351)	(4.576)	(8.095)	(14.709)
Mean R^2	0.296	0.317	0.336	0.379	0.385	0.343	0.333
Median R^2	0.311	0.339	0.349	0.397	0.412	0.362	0.355
Obs.	80380	120176	94332	54815	31690	54209	435602
Bonds	1259	1857	1257	850	502	951	6676

Table III
Principal Component Analysis.

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \varepsilon_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section II, with $\Delta Lev_{i,t}$ as the change in firm leverage, ΔRF_t the change in 10-year Treasury interest rate, ΔRF_t^2 the squared change in the 10-year Treasury interest rate, $\Delta Slope_t$ the change in the slope of the term structure, ΔVIX_t the change in ΔVIX_t index, RM_t the S&P 500 return, and $\Delta Jump_t$ the change in a jump factor based on S&P 500 index options. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. For each bin, we report the number of bonds, the number of observations, the principal components loadings and the ratio of variation of the residual to the total variation. We further report the proportions of the variance of the residuals explained by the first and second principal components, PC1 and PC2, respectively and the total unexplained variance of the regression in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

Leverage	Maturity	Bonds	Observations	PC1	PC2	Exp
1	1	835	33695	0.098	0.104	0.012
1	2	599	14983	0.100	0.148	0.011
1	3	643	25188	0.069	0.138	0.007
2	1	1207	46064	0.122	0.094	0.016
2	2	949	23159	0.136	0.161	0.018
2	3	1046	41176	0.094	0.150	0.010
3	1	728	27117	0.195	0.157	0.039
3	2	652	16004	0.176	0.217	0.030
3	3	664	33202	0.144	0.182	0.021
4	1	473	16155	0.284	0.162	0.078
4	2	455	11210	0.249	0.258	0.059
4	3	398	15351	0.162	0.222	0.030
5	1	311	10239	0.307	-0.297	0.099
5	2	286	7360	0.323	0.143	0.099
5	3	204	7557	0.206	0.247	0.052
6	1	697	22493	0.396	-0.534	0.150
6	2	608	14755	0.385	-0.417	0.141
6	3	334	11865	0.359	0.120	0.130
Proportion of Variance				0.802	0.042	
Unexplained Variance					1.405	

Table IV
Correlation Matrix of Changes in Yield Spreads and Proxies of Inflation Risk

This table reports the standard deviation and correlation matrix of the changes in yield spreads ($\Delta YS_{i,t}$), the changes in the proxies of inflation risk ($\Delta Swap_{i,t}$, $\Delta Swap^2_{i,t}$, $\Delta\sigma^{(21)}_{i,t}$, $\Delta\sigma^{(120)}_{i,t}$, $\Delta Slope^S_t$) introduced in Section IV. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	SD	$\Delta YS_{i,t}$	$\Delta Swap_{i,t}$	$\Delta Swap^2_{i,t}$	$\Delta\sigma^{(21)}_{i,t}$	$\Delta\sigma^{(120)}_{i,t}$	$\Delta Slope^S_t$
$\Delta YS_{i,t}$	0.856	1	-0.281	0.233	0.229	0.209	0.248
$\Delta Swap_{i,t}$	0.216		1	-0.320	-0.277	-0.365	-0.586
$\Delta Swap^2_{i,t}$	0.123			1	0.443	0.382	0.322
$\Delta\sigma^{(21)}_{i,t}$	0.061				1	0.182	0.329
$\Delta\sigma^{(120)}_{i,t}$	0.066					1	0.382
$\Delta Slope^S_t$	0.254						1

Table V
Inflation Risk and Yield Spread Changes

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t}, \Delta RF_t, \Delta RF_t^2, \Delta Slope_t, \Delta VIX_t, RM_t, \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t}, \Delta Swap^2_{i,t}, \Delta\sigma^{(21)}_{i,t}, \Delta\sigma^{(120)}_{i,t}, \Delta Slope^S_t]$ refers to the proxies for inflation risk introduced in Section IV. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. We report the fraction of variance explained as defined in Eq.6, the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the total unexplained variance in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Individual Bond Regressions							
Intercept	0.040	0.036	-0.005	0.033	0.047	0.043	0.019
	(19.514)	(18.045)	(-2.279)	(17.484)	(23.031)	(21.268)	(8.991)
$\Delta Lev_{i,t}$	0.020	0.015	0.017	0.024	0.016	0.013	0.013
	(17.671)	(13.276)	(15.131)	(21.572)	(14.517)	(11.596)	(14.240)
ΔRF_t	-0.632	-0.449	-0.412	-0.548	-0.529	-0.581	-0.331
	(-43.549)	(-35.101)	(-32.695)	(-41.466)	(-40.948)	(-41.962)	(-25.155)
ΔRF_t^2	-0.009	-0.111	0.069	-0.046	-0.160	-0.077	-0.128
	(-0.345)	(-4.167)	(2.720)	(-1.874)	(-5.977)	(-2.982)	(-4.853)
$\Delta Slope_t$	0.565	0.535	0.193	0.420	0.394	0.472	0.214
	(27.638)	(28.233)	(10.990)	(22.888)	(21.667)	(24.549)	(12.229)
ΔVIX_t	0.006	0.007	0.006	0.005	0.006	0.003	0.004
	(7.927)	(8.029)	(7.578)	(5.768)	(6.706)	(2.993)	(4.846)
RM_t	-0.048	-0.034	-0.042	-0.040	-0.046	-0.043	-0.030
	(-42.538)	(-33.154)	(-39.726)	(-38.815)	(-40.721)	(-39.364)	(-27.274)
$\Delta Jump_t$	0.008	0.005	0.005	0.006	0.007	0.005	0.003
	(14.709)	(9.632)	(8.852)	(10.704)	(12.537)	(9.891)	(5.243)
$\Delta Swap_{i,t}$		-0.667					-0.383
		(-36.930)					(-21.121)
$\Delta Swap_{i,t}^2$			1.085				0.411
			(23.158)				(8.488)
$\Delta\sigma_{i,t}^{(21)}$				1.822			1.003
				(34.356)			(21.349)
$\Delta\sigma_{i,t}^{(120)}$					1.280		0.484
					(28.376)		(12.060)
$\Delta Slope_t^S$						0.480	0.161
						(33.270)	(10.961)
Mean R^2	0.333	0.365	0.386	0.397	0.368	0.366	0.461
Median R^2	0.355	0.388	0.402	0.407	0.388	0.392	0.479
Obs.	435602	435602	435602	435602	435602	435602	435602
Bonds	6676	6676	6676	6676	6676	6676	6676
Panel B: Principal Component Analysis							
FVE	0	0.184	0.241	0.211	0.179	0.178	0.499
PC1	0.802	0.753	0.744	0.756	0.767	0.768	0.687
PC2	0.042	0.070	0.059	0.056	0.058	0.051	0.090
UV	1.405	1.147	1.067	1.109	1.154	1.154	0.703

Table VI
Inflation Risk and Yield Spread Changes: Group by Leverage

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV. We assign each bond to a cohort based on the firm's average leverage ratio (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the number of observations and bonds in the sample, respectively. We also report the results across all bonds. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)
	<15%	15%–25%	25%–35%	35%–45%	45%–55%	>55%
Intercept	0.005	0.003	0.012	0.012	0.021	0.083
	(1.583)	(1.395)	(3.967)	(2.505)	(2.429)	(7.438)
$\Delta Lev_{i,t}$	0.007	0.003	0.010	0.015	0.024	0.041
	(2.288)	(2.492)	(7.846)	(10.138)	(6.443)	(10.383)
ΔRF_t	−0.209	−0.236	−0.299	−0.423	−0.333	−0.635
	(−16.239)	(−19.669)	(−13.053)	(−13.233)	(−5.537)	(−9.162)
ΔRF_t^2	0.007	−0.008	−0.139	−0.157	−0.282	−0.420
	(0.128)	(−0.331)	(−3.188)	(−2.443)	(−2.916)	(−3.208)
$\Delta Slope_t$	0.182	0.173	0.195	0.277	0.184	0.320
	(8.163)	(10.337)	(7.074)	(7.110)	(2.728)	(3.315)
ΔVIX_t	0.006	0.004	0.006	0.006	0.010	−0.003
	(2.261)	(5.789)	(4.491)	(3.570)	(2.771)	(−0.834)
RM_t	−0.005	−0.012	−0.024	−0.034	−0.052	−0.089
	(−2.026)	(−14.207)	(−16.109)	(−15.163)	(−12.221)	(−17.474)
$\Delta Jump_t$	0.001	0.001	0.002	0.004	0.003	0.009
	(1.074)	(1.393)	(2.258)	(3.230)	(1.409)	(3.235)
$\Delta Swap_{i,t}$	−0.131	−0.187	−0.281	−0.452	−0.779	−0.962
	(−5.234)	(−11.575)	(−9.231)	(−9.497)	(−10.675)	(−10.389)
$\Delta Swap_{i,t}^2$	0.084	0.244	0.415	0.443	1.095	0.775
	(1.250)	(5.537)	(5.510)	(2.790)	(4.047)	(3.493)
$\Delta\sigma_{i,t}^{(21)}$	0.644	0.899	1.046	1.561	0.966	1.147
	(11.548)	(17.590)	(11.433)	(11.500)	(4.096)	(5.225)
$\Delta\sigma_{i,t}^{(120)}$	0.114	0.335	0.370	0.603	0.299	1.410
	(1.574)	(8.747)	(6.499)	(7.032)	(1.549)	(6.906)
$\Delta Slope_t^S$	0.078	0.122	0.143	0.204	0.070	0.377
	(3.622)	(9.191)	(6.009)	(5.717)	(1.147)	(4.873)
Mean R^2	0.415	0.455	0.458	0.520	0.501	0.466
Median R^2	0.423	0.463	0.465	0.558	0.520	0.500
Obs.	80380	120176	94332	54815	31690	54209
Bonds	1259	1857	1257	850	502	951

Table VII

Inflation Risk and Yield Spread Changes: Group by Rating

We assign each bond to 5 cohorts based on the bond rating (AAA-AA, A, BBB, BB and B-C), and for each industrial bond i with at least 25 monthly observations of yield spread changes $\Delta YS_{i,t}$ within the cohort, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV. We report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. We also report the results across all bonds. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	AAA-AA		A		BBB		BB		B-C	
Intercept	0.011	0.007	0.010	0.007	0.025	0.005	0.057	0.043	0.095	0.088
	(3.897)	(2.087)	(6.123)	(4.004)	(12.404)	(2.591)	(9.961)	(7.346)	(7.545)	(6.706)
$\Delta Lev_{i,t}$	0.004	0.004	0.011	0.005	0.012	0.008	0.020	0.010	0.060	0.044
	(1.503)	(1.326)	(3.921)	(3.514)	(14.531)	(8.944)	(8.401)	(5.064)	(12.586)	(9.702)
ΔRF_t	−0.250	−0.207	−0.305	−0.204	−0.480	−0.187	−1.108	−0.687	−1.214	−0.656
	(−22.676)	(−14.901)	(−34.802)	(−21.038)	(−36.319)	(−17.799)	(−24.430)	(−16.999)	(−12.758)	(−7.127)
ΔRF_t^2	0.005	−0.019	0.112	0.051	0.022	−0.029	−0.112	−0.217	0.076	−0.317
	(0.090)	(−0.285)	(4.379)	(2.100)	(0.762)	(−1.047)	(−1.586)	(−3.049)	(0.440)	(−1.945)
$\Delta Slope_t$	0.211	0.152	0.267	0.157	0.488	0.096	0.938	0.467	0.544	0.136
	(10.490)	(7.902)	(18.003)	(11.069)	(22.586)	(6.662)	(13.544)	(7.096)	(3.960)	(1.092)
ΔVIX_t	0.005	0.005	0.007	0.005	0.001	0.002	0.011	0.009	0.013	−0.004
	(2.638)	(3.095)	(9.293)	(7.566)	(1.197)	(1.774)	(4.742)	(2.658)	(2.945)	(−0.795)
RM_t	−0.009	−0.005	−0.015	−0.009	−0.038	−0.019	−0.066	−0.045	−0.128	−0.108
	(−4.587)	(−3.331)	(−21.279)	(−13.407)	(−33.458)	(−22.400)	(−19.892)	(−11.405)	(−22.447)	(−18.338)
$\Delta Jump_t$	0.001	0.000	0.003	0.001	0.005	0.000	0.014	0.009	0.024	0.014
	(0.910)	(−0.091)	(5.971)	(2.646)	(8.894)	(0.962)	(8.754)	(5.706)	(7.791)	(4.286)
$\Delta Swap_{i,t}$		−0.045		−0.141		−0.312		−0.516		−1.187
		(−1.686)		(−13.173)		(−18.142)		(−8.419)		(−10.517)
$\Delta Swap_{i,t}^2$		0.096		0.087		0.237		0.365		0.340
		(2.394)		(2.624)		(4.767)		(2.142)		(0.948)
$\Delta\sigma_{i,t}^{(21)}$		0.378		0.634		1.161		1.121		0.896
		(9.453)		(17.913)		(22.496)		(8.481)		(3.458)
$\Delta\sigma_{i,t}^{(120)}$		−0.024		0.163		0.382		0.178		1.053
		(−0.692)		(6.909)		(9.757)		(1.407)		(4.298)
$\Delta Slope_t^S$		0.024		0.043		0.111		0.099		0.140
		(0.833)		(4.077)		(8.771)		(2.456)		(1.622)
Mean R^2	0.275	0.348	0.307	0.411	0.328	0.485	0.420	0.513	0.350	0.436
Median R^2	0.287	0.348	0.320	0.410	0.356	0.502	0.451	0.552	0.366	0.464
Obs.	34861	34861	113562	113562	166412	166412	46814	46814	41083	41083
Bonds	538	538	1825	1825	2791	2791	937	937	850	850

Table VIII
Inflation Risk and Yield Spread Changes: Group by Time To Maturity

We assign each bond to 3 cohorts based on the bond time to maturity (less than five years, five to twelve years, and over twelve years), and for each industrial bond i with at least 25 monthly observations of yield spread changes $\Delta YS_{i,t}$ within the cohort, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV. We report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. We also report the results across all bonds. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)
	<5 Years		5-12 Years		>12 Years	
Intercept	0.034	0.007	0.030	0.025	0.026	0.017
	(9.932)	(1.965)	(7.495)	(5.988)	(11.586)	(8.663)
$\Delta Lev_{i,t}$	0.020	0.014	0.024	0.018	0.010	0.005
	(10.422)	(9.714)	(4.444)	(3.589)	(7.598)	(3.230)
ΔRF_t	-0.600	-0.245	-0.541	-0.362	-0.544	-0.294
	(-22.395)	(-10.782)	(-16.240)	(-10.287)	(-32.047)	(-23.840)
ΔRF_t^2	-0.059	0.033	0.071	0.049	0.067	-0.216
	(-1.253)	(0.739)	(1.524)	(1.013)	(2.435)	(-7.556)
$\Delta Slope_t$	0.704	0.255	0.287	0.088	0.266	0.004
	(17.866)	(8.238)	(5.386)	(1.987)	(12.467)	(0.316)
ΔVIX_t	0.007	0.008	-0.001	-0.004	0.004	0.002
	(5.888)	(5.486)	(-1.143)	(-3.373)	(5.217)	(2.846)
RM_t	-0.050	-0.028	-0.060	-0.044	-0.034	-0.023
	(-26.782)	(-15.343)	(-27.064)	(-20.282)	(-28.070)	(-24.896)
$\Delta Jump_t$	0.011	0.003	0.012	0.007	0.003	0.000
	(10.951)	(3.822)	(11.958)	(6.361)	(5.366)	(0.047)
$\Delta Swap_{i,t}$		-0.357		-0.407		-0.384
		(-12.994)		(-11.588)		(-15.574)
$\Delta Swap_{i,t}^2$		0.022		-0.178		0.890
		(0.383)		(-1.051)		(10.420)
$\Delta\sigma_{i,t}^{(21)}$		0.688		1.023		1.077
		(10.615)		(10.570)		(18.781)
$\Delta\sigma_{i,t}^{(120)}$		0.121		0.192		0.642
		(2.315)		(2.311)		(12.750)
$\Delta Slope_t^S$		0.056		0.133		0.066
		(2.089)		(6.023)		(7.247)
Mean R^2	0.263	0.396	0.364	0.444	0.396	0.495
Median R^2	0.261	0.380	0.389	0.452	0.403	0.497
Obs.	160567	160567	72570	72570	117393	117393
Bonds	3598	3598	2167	2167	1564	1564

Table IX
Inflation Risk and Yield Spread Changes: Non-linearities

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} \times H_{t-1} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV, and H_{t-1} is a dummy variable taking the value of 1 when the 1-year inflation swap rate in month $t-1$ is above the 80% percentile, we define these months as "high" inflation months. We further restrict the sample to Bondswith at least 6 monthly observations in "high" inflation months and likewise outside. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. We report the fraction of variance explained as defined in Eq.6, the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the total unexplained variance in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Intercept	0.038	0.031	-0.021	0.030	0.049	0.043	0.016	0.011
	(14.267)	(12.060)	(-7.313)	(12.692)	(17.784)	(16.380)	(5.552)	(3.471)
$\Delta Lev_{i,t}$	0.024	0.016	0.018	0.029	0.017	0.014	0.015	0.015
	(13.675)	(9.981)	(10.665)	(16.776)	(10.489)	(8.505)	(10.529)	(9.142)
ΔRF_t	-0.664	-0.398	-0.382	-0.551	-0.520	-0.574	-0.301	-0.297
	(-38.705)	(-30.924)	(-28.782)	(-38.029)	(-37.883)	(-38.986)	(-24.566)	(-21.815)
ΔRF_t^2	-0.108	-0.314	-0.051	-0.136	-0.358	-0.202	-0.321	-0.287
	(-2.839)	(-7.738)	(-1.387)	(-3.770)	(-8.987)	(-5.286)	(-8.335)	(-6.545)
$\Delta Slope_t$	0.608	0.584	0.152	0.415	0.389	0.504	0.157	0.183
	(30.362)	(30.873)	(8.851)	(24.417)	(24.470)	(27.609)	(9.954)	(10.524)
ΔVIX_t	0.012	0.011	0.012	0.008	0.010	0.005	0.009	0.007
	(9.927)	(9.184)	(9.188)	(6.333)	(7.950)	(3.904)	(6.317)	(3.949)
RM_t	-0.044	-0.025	-0.036	-0.035	-0.042	-0.040	-0.025	-0.025
	(-29.551)	(-18.155)	(-25.553)	(-25.684)	(-27.084)	(-27.321)	(-17.017)	(-14.732)
$\Delta Jump_t$	0.010	0.006	0.005	0.006	0.007	0.005	0.003	0.003
	(11.761)	(7.828)	(6.164)	(7.869)	(9.257)	(6.288)	(3.503)	(3.626)
$\Delta Swap_{i,t}$		-0.933					-0.309	-0.385
		(-32.336)					(-14.132)	(-12.217)
$\Delta Swap_{i,t} \times H$		1.069						0.479
		(32.394)						(10.416)
$\Delta Swap_{i,t}^2$			1.567				0.690	0.667
			(27.361)				(14.495)	(10.437)
$\Delta Swap_{i,t}^2 \times H$			0.455					0.272
			(6.323)					(1.941)
$\Delta\sigma_{i,t}^{(21)}$				2.419			1.121	1.230
				(28.249)			(20.567)	(17.636)
$\Delta\sigma_{i,t}^{(21)} \times H$				-1.889				-1.830
				(-20.201)				(-15.830)
$\Delta\sigma_{i,t}^{(120)} \times H$					-1.666			-0.574
					(-20.957)			(-5.168)
$\Delta\sigma_{i,t}^{(120)}$					2.098		0.841	0.842
					(31.268)		(17.255)	(9.865)
$\Delta Slope_t^S$						0.624	0.191	0.188
						(29.599)	(10.553)	(7.787)
$\Delta Slope_t^S \times H$						-0.821		-0.169
						(-30.388)		(-3.969)
Mean R^2	0.376	0.437	0.446	0.469	0.427	0.436	0.538	0.540
Median R^2	0.396	0.461	0.469	0.479	0.450	0.462	0.555	0.563
Obs.	267290	267290	267290	267290	267290	267290	267290	267290
Bonds	3749	3749	3749	3749	3749	3749	3749	3749
Panel B: Principal Component Analysis								
FVE		0.226	0.232	0.220	0.197	0.204	0.435	0.489
PC1	0.713	0.626	0.631	0.639	0.655	0.651	0.554	0.531
PC2	0.067	0.095	0.089	0.084	0.083	0.080	0.119	0.127
UV	1.621	1.255	1.244	1.264	1.301	1.290	0.916	0.828

Table X
Different Residual Groups

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the CDGM baseline models:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \varepsilon_{i,t},$$

and the model including inflation risk proxies:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV. We then assign each month's residuals to the intersection of two cohorts based on either three maturity groups (less than five years, five to twelve years, and over twelve years), six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%), five rating groups (AAA-AA, A, BBB, BB and B-C), five volume groups based on the total dollar volume traded in the previous month, five market beta groups and five VIX beta based on their regression betas on the SP 500 and the ΔVIX_t in Eq. 2. After, we assign the regression residuals to these 18 or 15 cohorts depending on the groups, we compute an average residual. We then extract the principal components of the covariance matrix of these residuals. For each pair of grouping variables, we report the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the fraction of variance explained as defined in Eq. 6. The first row reports the results of the baseline model, while the second row reports results of the model including inflation risk proxies. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

Panel A: Baseline

Group 1	Group 2	Inflation Risk	PC1	PC2	FVE
Time to Maturity	Leverage	No	0.802	0.042	
		Yes	0.687	0.090	0.499
Time to Maturity	Rating	No	0.785	0.064	
		Yes	0.661	0.106	0.509
Time to Maturity	Volume	No	0.820	0.062	
		Yes	0.737	0.089	0.502
Time to Maturity	Market Beta	No	0.767	0.081	
		Yes	0.643	0.102	0.519
Time to Maturity	VIX Beta	No	0.763	0.071	
		Yes	0.620	0.106	0.554

Panel B: Non-linear effects

Group 1	Group 2	Inflation Risk	PC1	PC2	FVE
Time to Maturity	Leverage	No	0.713	0.067	
		Yes	0.531	0.127	0.489
Time to Maturity	Rating	No	0.688	0.091	
		Yes	0.518	0.157	0.473
Time to Maturity	Volume	No	0.722	0.132	
		Yes	0.615	0.171	0.461
Time to Maturity	Market Beta	No	0.764	0.075	
		Yes	0.613	0.119	0.539
Time to Maturity	VIX Beta	No	0.714	0.071	
		Yes	0.520	0.119	0.560

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Internet Appendix for: ”Inflation Risk and Yield Spread Changes”

Diego Bonelli

This Internet Appendix contains supplemental material for the article ”Inflation Risk and Yield Spread Changes”

- Table IA.I presents time-series regressions of yield spread changes onto inflation proxies where average coefficients and their statistical significance are computed within industry PPI groups. We compute the average absolute variation of the industry Producer Price Index (PPI), during the life of each bond. We assign each bond to a cohort based on the bond's average absolute PPI and report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The PPI data comes from the U.S. Bureau of Labor Statistics.

- Table IA.II presents time-series regressions of yield spread changes onto inflation proxies based on off-the-run TIPS rates.

- Table IA.III presents time-series regressions of yield spread changes onto inflation proxies controlling for change in cash to market value, Ceballos (2021)'s inflation volatility risk (IVR) and changes in unemployment and real consumption and income. Inflation, unemployment and real consumption and income data comes from the Federal Reserve Bank of St. Louis.

- Table IA.IV presents time-series regressions of yield spread changes onto inflation proxies controlling for Friewald and Nagler (2019) and He et al. (2022) variables.

Table IA.I
Inflation Risk and Yield Spread Changes: Group by PPI

We assign each bond to 5 cohorts based on the average absolute variation of the industry Producer Price Index (PPI), during the life of each bond, and for each industrial bond i with at least 25 monthly observations of yield spread changes $\Delta YS_{i,t}$ within the cohort, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap^2_{i,t},\ \Delta\sigma^{(21)}_{i,t},\ \Delta\sigma^{(120)}_{i,t},\ \Delta Slope^S_t]$ refers to the proxies for inflation risk introduced in Section IV. We report the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. We also report the results across all bonds. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Low		2		3		4		High	
Intercept	0.023	0.007	0.020	−0.002	0.033	0.021	0.042	0.034	0.080	0.022
	(6.552)	(2.095)	(5.677)	(−0.695)	(7.312)	(4.701)	(8.139)	(6.567)	(11.096)	(3.142)
$\Delta Lev_{i,t}$	0.010	0.006	0.017	0.008	0.027	0.019	0.021	0.018	0.023	0.016
	(6.143)	(4.179)	(5.462)	(4.334)	(6.177)	(5.495)	(8.942)	(6.891)	(9.526)	(7.271)
ΔRF_t	−0.515	−0.326	−0.489	−0.240	−0.521	−0.250	−0.508	−0.331	−1.049	−0.330
	(−20.670)	(−11.289)	(−25.921)	(−13.252)	(−13.238)	(−6.854)	(−16.271)	(−10.829)	(−18.135)	(−7.719)
ΔRF_t^2	0.089	0.003	0.116	−0.001	0.037	−0.070	−0.065	−0.218	−0.127	−0.261
	(2.013)	(0.052)	(2.051)	(−0.011)	(0.728)	(−1.485)	(−1.113)	(−3.271)	(−1.449)	(−3.132)
$\Delta Slope_t$	0.504	0.273	0.451	0.123	0.466	0.151	0.418	0.222	0.781	0.096
	(14.255)	(6.599)	(16.080)	(4.513)	(9.183)	(3.480)	(9.885)	(5.923)	(8.874)	(1.780)
ΔVIX_t	0.001	−0.001	0.005	0.006	0.007	0.003	0.010	0.005	0.004	0.004
	(0.545)	(−0.474)	(3.763)	(4.379)	(4.483)	(1.781)	(4.759)	(2.562)	(1.850)	(1.926)
RM_t	−0.036	−0.025	−0.033	−0.015	−0.042	−0.028	−0.045	−0.038	−0.083	−0.038
	(−17.010)	(−11.463)	(−19.832)	(−8.619)	(−15.985)	(−10.988)	(−17.204)	(−14.104)	(−21.098)	(−12.595)
$\Delta Jump_t$	0.006	0.002	0.006	0.001	0.005	0.000	0.008	0.004	0.013	0.002
	(5.504)	(2.007)	(5.520)	(1.125)	(4.128)	(0.343)	(5.898)	(2.912)	(7.518)	(1.216)
$\Delta Swap_{i,t}$		−0.207		−0.244		−0.377		−0.271		−1.022
		(−7.413)		(−8.589)		(−8.193)		(−6.110)		(−14.798)
$\Delta Swap^2_{i,t}$		0.373		0.441		0.239		0.329		0.797
		(4.480)		(7.351)		(2.294)		(3.328)		(3.698)
$\Delta\sigma^{(21)}_{i,t}$		0.670		0.912		0.836		0.721		2.201
		(8.364)		(11.732)		(6.354)		(7.714)		(12.802)
$\Delta\sigma^{(120)}_{i,t}$		0.519		0.408		0.391		0.519		0.462
		(7.046)		(7.555)		(4.745)		(5.873)		(3.029)
$\Delta Slope^S_t$		0.158		0.164		0.098		0.107		0.239
		(6.054)		(6.008)		(2.998)		(2.315)		(4.521)
Mean R^2	0.320	0.446	0.355	0.495	0.311	0.426	0.328	0.440	0.345	0.504
Median R^2	0.345	0.469	0.387	0.528	0.324	0.420	0.351	0.461	0.366	0.544
Obs.	67215	67215	67036	67036	67412	67412	66792	66792	67031	67031
Bonds	1112	1112	1172	1172	976	976	1003	1003	1053	1053

Table IA.II
Inflation Risk and Yield Spread Changes: TIPS

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta TIPS_{i,t},\ \Delta TIPS_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)T},\ \Delta\sigma_{i,t}^{(120)T},\ \Delta Slope_t^T]$ refers to the proxies for inflation risk computed using TIPS' rates following Section IV. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. We report the fraction of variance explained as defined in Eq.6, the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the total unexplained variance in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Individual Bond Regressions							
Intercept	0.040	0.028	-0.002	0.036	0.031	0.050	0.018
	(19.514)	(14.765)	(-0.850)	(18.402)	(16.653)	(24.542)	(9.044)
$\Delta Lev_{i,t}$	0.020	0.013	0.016	0.012	0.023	0.016	0.013
	(17.671)	(11.779)	(13.675)	(10.943)	(22.420)	(14.757)	(13.442)
ΔRF_t	-0.632	-0.391	-0.375	-0.552	-0.545	-0.546	-0.310
	(-43.549)	(-31.305)	(-29.230)	(-41.339)	(-41.060)	(-42.120)	(-23.506)
ΔRF_t^2	-0.009	-0.040	0.088	0.005	0.016	-0.228	-0.040
	(-0.345)	(-1.544)	(3.402)	(0.191)	(0.686)	(-8.163)	(-1.505)
$\Delta Slope_t$	0.565	0.470	0.087	0.360	0.369	0.350	0.176
	(27.638)	(26.306)	(5.008)	(19.673)	(20.976)	(19.901)	(10.255)
ΔVIX_t	0.006	0.006	0.006	0.003	0.004	0.006	0.004
	(7.927)	(7.003)	(7.631)	(3.556)	(4.373)	(7.116)	(5.900)
RM_t	-0.048	-0.029	-0.041	-0.041	-0.040	-0.045	-0.028
	(-42.538)	(-28.578)	(-39.746)	(-38.477)	(-38.731)	(-41.718)	(-27.948)
$\Delta Jump_t$	0.008	0.004	0.003	0.004	0.005	0.005	0.001
	(14.709)	(7.391)	(6.193)	(6.749)	(9.800)	(9.813)	(2.085)
$\Delta TIPS_{i,t}$		-0.698					-0.470
		(-41.044)					(-26.636)
$\Delta TIPS_{i,t}^2$			0.843				0.208
			(20.618)				(5.582)
$\Delta Slope_t^T$				0.457			0.016
				(35.534)			(1.311)
$\Delta \sigma_{i,t}^{(21)T}$					1.607		0.907
					(33.161)		(22.011)
$\Delta \sigma_{i,t}^{(120)T}$						1.194	0.284
						(28.022)	(7.801)
Mean R^2	0.333	0.378	0.396	0.373	0.405	0.367	0.465
Median R^2	0.355	0.406	0.410	0.400	0.411	0.388	0.481
Obs.	435602	435602	435602	435602	435602	435602	435602
Bonds	6676	6676	6676	6676	6676	6676	6676
Panel B: Principal Component Analysis							
FVE	0	0.308	0.286	0.234	0.209	0.207	0.549
PC1	0.802	0.715	0.738	0.751	0.768	0.765	0.686
PC2	0.042	0.084	0.060	0.055	0.052	0.063	0.086
UV	1.405	0.972	1.003	1.076	1.112	1.115	0.634

Table IA.III

Inflation Risk and Yield Spread Changes: Other Variables

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \boldsymbol{\Gamma}_i^T \boldsymbol{\Delta C}_{i,t} + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV and the vector $\boldsymbol{\Delta C}_{i,t} := [\Delta Cash/ME_{i,t},\ \Delta IVR_t,\ \Delta Consumption_t,\ \Delta Income_t,\ \Delta Unemployment_t]$ refers to the changes in control proxies of cash over market value, inflation volatility risk, real consumption, real income and unemployment. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. We report the fraction of variance explained as defined in Eq.6, the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the total unexplained variance in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Individual Bond Regressions						
Intercept	0.039	0.019	0.078	0.047	0.026	0.018
	(18.947)	(8.770)	(23.489)	(13.834)	(11.032)	(6.456)
$\Delta Lev_{i,t}$	0.016	0.011	0.020	0.013	0.017	0.013
	(13.515)	(9.955)	(17.333)	(13.164)	(14.581)	(12.718)
ΔRF_t	-0.630	-0.334	-0.636	-0.337	-0.473	-0.336
	(-43.131)	(-25.058)	(-43.844)	(-25.567)	(-37.271)	(-25.380)
ΔRF_t^2	-0.010	-0.117	-0.007	-0.119	0.173	-0.050
	(-0.386)	(-4.003)	(-0.278)	(-4.434)	(6.621)	(-1.724)
$\Delta Slope_t$	0.569	0.219	0.590	0.233	0.347	0.212
	(27.997)	(12.731)	(28.129)	(12.825)	(19.785)	(11.741)
ΔVIX_t	0.006	0.004	0.006	0.004	0.003	0.001
	(8.007)	(4.956)	(7.690)	(4.702)	(3.536)	(0.691)
RM_t	-0.047	-0.030	-0.048	-0.030	-0.041	-0.034
	(-42.372)	(-26.740)	(-42.145)	(-27.218)	(-39.742)	(-28.726)
$\Delta Jump_t$	0.008	0.003	0.008	0.003	0.004	0.002
	(15.234)	(5.897)	(14.569)	(4.989)	(6.936)	(4.031)
$\Delta Cash/ME_{i,t}$	6.151	6.573				
	(2.366)	(2.385)				
ΔIVR_t			-0.164	-0.149		
			(-13.467)	(-11.569)		
$\Delta Consumption_t$					-4.469	-3.015
					(-7.247)	(-4.980)
$\Delta Income_t$					3.617	4.686
					(9.083)	(9.813)
$\Delta Unemployment_t$					-0.036	-0.040
					(-3.221)	(-3.787)
$\Delta Swap_{i,t}$		-0.371		-0.378		-0.282
		(-19.917)		(-20.697)		(-16.185)
$\Delta Swap_{i,t}^2$		0.384		0.531		0.248
		(7.811)		(10.646)		(4.844)
$\Delta\sigma_{i,t}^{(21)}$		1.007		0.917		0.632
		(21.329)		(19.321)		(13.388)
$\Delta\sigma_{i,t}^{(120)}$		0.482		0.398		0.525
		(12.012)		(9.491)		(12.622)
$\Delta Slope_t^S$		0.168		0.156		0.118
		(11.278)		(10.688)		(8.119)
Mean R2	0.338	0.465	0.330	0.459	0.425	0.487
Median R2	0.360	0.483	0.356	0.478	0.425	0.506
obs	435602	435602	435602	435602	435602	435602
bonds	6676	6676	6676	6676	6676	6676
Mean R^2	0.349	0.462	0.336	0.454	0.423	0.481
Median R^2	0.371	0.478	0.362	0.472	0.423	0.499
Obs.	377573	377573	377573	377573	377573	377573
Bonds	5905	5905	5905	5905	5905	5905
Panel B: Principal Component Analysis						
FVE	0.051	0.529	0.024	0.514	0.305	0.569
PC1	0.808	0.689	0.800	0.685	0.753	0.685
PC2	0.040	0.092	0.043	0.091	0.055	0.088
UV	1.334	0.662	1.372	0.682	0.977	0.606

Table IA.IV

Inflation Risk and Yield Spread Changes: FN and HKS

For each industrial bond i with at least 25 monthly observations of yield spread changes, $\Delta YS_{i,t}$, we estimate the model:

$$\Delta YS_{i,t} = \alpha_i + \boldsymbol{\beta}_i^T \boldsymbol{\Delta S}_{i,t} + \boldsymbol{\theta}_i^T \boldsymbol{\Delta I}_{i,t} + \boldsymbol{\Gamma}_i^T \boldsymbol{\Delta C}_t + \nu_{i,t},$$

where $\boldsymbol{\Delta S}_{i,t} := [\Delta Lev_{i,t},\ \Delta RF_t,\ \Delta RF_t^2,\ \Delta Slope_t,\ \Delta VIX_t,\ RM_t,\ \Delta Jump_t]$ is the vector of the structural model variables defined in Section III. The vector $\boldsymbol{\Delta I}_{i,t} := [\Delta Swap_{i,t},\ \Delta Swap_{i,t}^2,\ \Delta\sigma_{i,t}^{(21)},\ \Delta\sigma_{i,t}^{(120)},\ \Delta Slope_t^S]$ refers to the proxies for inflation risk introduced in Section IV and the vector $\boldsymbol{\Delta C}_{i,t} := [\Delta Inv_t,\ \Delta amt_t,\ \Delta block.trd_t,\ \Delta match.trd_t,\ \Delta ig2junk_t,\ \Delta ted_t,\ \Delta Distress_t,\ ]$ refers to the changes in FN and HKS proxies. Panel A reports the average coefficients across bonds, the associated t-statistics, the mean and median adjusted R2 values, and the numbers of observations and bonds in the sample, respectively. The t-statistics are calculated from the cross-sectional variation over the estimates for each coefficient. That is, we divide each reported coefficient value by the standard deviation of the estimates and scale by the square root of the number of bonds. Panel B reports the results of a principal component analysis on the residuals. We then assign each month's residuals to one of 18 bins defined by three maturity groups (less than five years, five to twelve years, and over twelve years) and six leverage groups (less than 15%, 15%–25%, 25%–35%, 35%–45%, 45%–55%, and greater than 55%) and compute an average residual. We extract the principal components of the covariance matrix of these residuals. We report the fraction of variance explained as defined in Eq.6, the proportions of variance explained by the first and second principal components, PC1 and PC2, respectively, and the total unexplained variance in percentage points. The sample is based on U.S. corporate bond transaction data from TRACE for the period 2005–2021.

	(1)	(2)	(3)	(4)
Panel A: Individual Bond Regressions				
Intercept	−0.024	−0.008	−0.031	−0.010
	(−6.408)	(−1.949)	(−7.928)	(−2.387)
$\Delta Lev_{i,t}$	0.019	0.014	0.018	0.013
	(15.198)	(10.616)	(15.317)	(10.532)
ΔRF_t	−0.421	−0.328	−0.398	−0.342
	(−33.390)	(−23.507)	(−30.412)	(−22.661)
ΔRF_t^2	0.234	−0.049	0.179	−0.074
	(8.838)	(−1.632)	(6.756)	(−2.394)
$\Delta Slope_t$	0.200	0.163	0.202	0.193
	(11.469)	(8.492)	(11.678)	(9.494)
ΔVIX_t	0.006	0.003	0.007	0.004
	(6.696)	(3.200)	(7.761)	(3.907)
RM_t	−0.038	−0.030	−0.029	−0.024
	(−35.984)	(−25.023)	(−26.976)	(−20.010)
$\Delta Jump_t$	0.003	0.002	0.003	0.002
	(5.503)	(2.689)	(5.195)	(2.714)
ΔInv_t	8.484	9.684	5.983	7.834
	(9.444)	(9.703)	(6.943)	(7.796)
Δamt_t	0.123	0.111	0.146	0.109
	(6.671)	(6.043)	(7.927)	(5.883)
$\Delta block.trd_t$	−0.509	−0.212	−0.609	−0.385
	(−3.901)	(−1.648)	(−4.691)	(−2.871)
$\Delta match.trd_t$	−0.271	−1.070	0.058	−0.991
	(−1.865)	(−6.564)	(0.390)	(−5.578)
$\Delta ig2junk_t$	0.001	0.002	0.001	0.001
	(5.959)	(7.435)	(5.586)	(6.466)
Δted_t	0.515	0.324	0.458	0.296
	(22.401)	(14.999)	(18.668)	(12.751)
$\Delta Distress_t$			0.067	0.037
			(19.087)	(9.760)
$\Delta Swap_{i,t}$		−0.324		−0.313
		(−16.057)		(−14.299)
$\Delta Swap_{i,t}^2$		0.283		0.342
		(5.340)		(6.277)
$\Delta \sigma_{i,t}^{(21)}$		0.642		0.613
		(13.441)		(11.231)
$\Delta \sigma_{i,t}^{(120)}$		0.529		0.449
		(11.869)		(10.012)
$\Delta Slope_t^S$		0.164		0.157
		(10.143)		(8.634)
Mean R2	0.438	0.500	0.449	0.508
Median R2	0.447	0.525	0.466	0.536
Obs.	435602	435602	435602	435602
Bonds	6676	6676	6676	6676
Panel B: Principal Component Analysis				
FVE	0.346	0.611	0.429	0.651
PC1	0.750	0.667	0.747	0.671
PC2	0.060	0.101	0.066	0.104
UV	0.919	0.547	0.803	0.491