

Credit Spreads and Business Cycle Fluctuations[†]

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Between the summer of 2007 and the spring of 2009, the US economy was gripped by an acute liquidity and credit crunch, by all accounts the most severe financial crisis since the Great Depression. Throughout this period of extreme financial turmoil, credit spreads—the difference in yields between various private debt instruments and government securities of comparable maturity—served as a crucial gauge of the degree of strains in the financial system. In addition, movements in credit spreads were thought to contain important signals regarding the evolution of the real economy and risks to the economic outlook, a view supported by the insights from the large literature on the predictive content of credit spreads—or asset prices more generally—for economic activity.¹

The focus on credit spreads is motivated, in part, by theories that depart from the paradigm of frictionless financial markets articulated by Modigliani and Miller (1958), theories that emphasize linkages between the quality of borrowers' balance sheets and their access to external finance. Fluctuations in credit spreads may also reflect shifts in the effective supply of funds offered by financial intermediaries, which, in the presence of financial market frictions, have important implications for the usefulness of credit spreads as predictors of economic activity. In the latter case, a deterioration in the capital position of financial intermediaries leads to a reduction in the supply of credit, causing an increase in the cost of debt finance—the widening of credit spreads—and a subsequent reduction in spending and production.

In this paper, we examine the relationship between corporate bond credit spreads and economic activity. To do so, we first construct a new credit spread index—the “GZ credit spread”—that has considerable predictive power for economic activity

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¹Financial indicators considered in this vast literature include stock prices (Fama 1981); spreads between long- and short-term risk-free interest rates (Harvey 1988, Estrella and Hardouvelis 1991; Estrella and Mishkin 1998; and Hamilton and Kim 2002); the Treasury term structure more generally (Ang, Piazzesi, and Wei 2006); spreads between rates on short-term commercial paper and rates on Treasury bills (Friedman and Kuttner 1992, 1998 and Emery 1996); and yield spreads on longer-term corporate bonds (Gertler and Lown 1999; Mueller 2009; Gilchrist, Yankov, and Zakrajšek 2009; and Faust et al. 2011). Stock and Watson (2003) provide a systematic evaluation of the information content of asset prices for short- and medium-term economic forecasting.

over the 1973–2010 period. Our approach builds on the recent work of Gilchrist, Yankov, and Zakrajšek (2009), in that we use prices of individual corporate bonds traded in the secondary market to construct this highly informative financial indicator. According to our results, the predictive ability of the GZ credit spread for future economic activity significantly exceeds that of the widely used default-risk indicators such as the standard Baa–Aaa corporate bond credit spread and the “paper-bill” spread.

As shown by Philippon (2009), the predictive content of corporate bond credit spreads for economic activity could reflect—absent any financial market frictions—the ability of the bond market to signal more accurately than the stock market a decline in economic fundamentals prior to a cyclical downturn. To address this issue, we use an empirical credit-spread pricing framework to decompose the GZ spread into two components: a component capturing the usual countercyclical movements in expected defaults, and a component representing the cyclical changes in the relationship between measured default risk and credit spreads—the so-called excess bond premium.

The decomposition is motivated in part by the existence of the “credit spread puzzle,” the well-known result from the corporate finance literature, showing that less than one-half of the variation in corporate bond credit spreads can be attributed to the financial health of the issuer; see, for example, Elton et al. (2001). As shown by Collin-Dufresne, Goldstein, and Martin (2001); Houweling, Mentink, and Vorst (2005); and Driessen (2005), the unexplained portion of the variation in credit spreads appears to reflect some combination of time-varying liquidity premium, the tax treatment of corporate bonds, and most importantly, a default-risk factor that captures compensation demanded by investors—above and beyond expected losses—for bearing exposure to corporate credit risk.

Our results indicate that a substantial portion of the information content of the GZ credit spread for economic activity can be attributed to deviations in the pricing of corporate bonds relative to the measured default risk of the issuer. We examine the macroeconomic implications of this finding using an identified vector autoregression (VAR) framework. According to our analysis, shocks to the excess bond premium that are orthogonal to the current state of the economy lead to economically and statistically significant declines in consumption, investment, and output, as well as to appreciable disinflation. Monetary policy is eased significantly in response to these adverse economic developments, and despite the decline in long-term Treasury yields, such shocks also imply a sharp fall in the stock market.

To provide an interpretation for these “financial disruptions” in the context of the 2007–2009 crisis, we examine how shocks to the profitability of primary dealers—the highly leveraged financial institutions that play a key role in the corporate cash market—affect credit supply conditions as measured by the excess bond premium. Our results indicate that an adverse shock to the equity valuations of these intermediaries—relative to the market return—leads to an immediate and persistent increase in their credit default swap (CDS) premiums, a response that is mirrored on an almost one-to-one basis by an increase in the excess bond premium.

The confluence of our results is thus consistent with the notion that an increase in the excess bond premium reflects a reduction in the effective risk-bearing capacity of the financial sector and, as a result, a contraction in the supply of credit; see, for

example, the theoretical work of He and Krishnamurthy (2010) and related empirical evidence of Adrian, Moench, and Shin (2010a, 2010b). Consistent with the financial accelerator mechanisms emphasized by Bernanke and Gertler (1989); Kiyotaki and Moore (1997); Bernanke, Gertler, and Gilchrist (1999); and Hall (2011), this reduction in credit availability augurs a change in financial conditions with significant adverse consequences for macroeconomic outcomes.

I. A High-Information Content Credit Spread Index

Academics, business economists, and policymakers have long relied on credit spreads to gauge the degree of strains in the financial system. In addition, market participants view these default-risk indicators as particularly useful for extracting investors' expectations of future economic outcomes, though obtaining an accurate reading of this information can be complicated by the presence of time-varying risk premiums. In this paper, we employ the "bottom-up" approach used by Gilchrist, Yankov, and Zakrajšek (2009) to construct a credit spread index with high information content for future economic developments. Importantly, we extend the time span of the analysis back to the mid-1970s, thereby covering an appreciably greater number of business cycles, a consideration of particular importance when one is evaluating the predictive ability of financial indicators for economic activity.

A. Data Sources and Methods

For a sample of US nonfinancial firms covered by the S&P's Compustat database and the Center for Research in Security Prices (CRSP), we obtained month-end secondary market prices of their outstanding securities from the Lehman/Warga and Merrill Lynch databases.² To ensure that we are measuring borrowing costs of different firms at the same point in their capital structure, we limited our sample to only senior unsecured issues with a fixed coupon schedule.

The micro-level aspect of our data allows us to construct credit spreads that are not subject to the "duration mismatch" that plagues the standard credit spread indexes. We do so by constructing a synthetic risk-free security that mimics exactly the cash flows of the corresponding corporate debt instrument. Specifically, consider a corporate bond k issued by firm i that at time t is promising a sequence of cash flows $\{C(s) : s = 1, 2, \dots, S\}$, consisting of the regular coupon payments and the repayment of the principle at maturity. The price of this bond is given by

$$P_{it}[k] = \sum_{s=1}^S C(s) D(t_s),$$

where $D(t) = e^{-r_t t}$ is the discount function in period t . To calculate the price of the corresponding risk-free security—denoted by $P_t^f[k]$ —we discount the cash-flow sequence $\{C(s) : s = 1, 2, \dots, S\}$ using continuously compounded zero-coupon Treasury yields in period t , obtained from the US Treasury yield curve estimated

²These two data sources include secondary market prices for a vast majority of dollar-denominated bonds publicly issued in the US corporate cash market; see Gilchrist, Yankov, and Zakrajšek (2009) for details.

TABLE 1—SUMMARY STATISTICS OF CORPORATE BOND CHARACTERISTICS

Variable	Mean	SD	Min	P50	Max
Number of bonds per firm/month	2.91	3.64	1.00	2.00	74.0
Market value of issue (\$mil.) ^a	322.9	326.6	1.22	238.6	5,628
Maturity at issue (years)	13.0	9.3	1.0	10.0	50.0
Term to maturity (years)	11.3	8.5	1.0	8.1	30.0
Duration (years)	6.47	3.20	0.91	6.06	16.0
Callable (pct.)	67.2	47.0	—	—	—
Credit rating (S&P)	—	—	D	BBB1	AAA
Coupon rate (pct.)	7.34	1.99	1.80	7.00	17.5
Nominal effective yield (pct.)	7.68	3.24	0.54	7.16	44.3
Credit spread (basis points)	204	281	5	118	3,499

Notes: Sample period: 1973:1–2010:9; Observations = 346,126; Number of bonds = 5,982; Number of firms = 1,112. Sample statistics are based on trimmed data (see text for details).

^aMarket value of the outstanding issue deflated by the CPI (2000 = 100).

by Gürkaynak, Sack, and Wright (2007). The resulting price $P_t^f[k]$ can then be used to calculate the yield—denoted by $y_t^f[k]$ —of a hypothetical Treasury security with exactly the same cash flows as the underlying corporate bond. The resulting credit spread $S_{it}[k] = y_{it}[k] - y_t^f[k]$, where $y_{it}[k]$ denotes the yield of the corporate bond k , is thus free of the bias that would occur were the spreads computed simply by matching the corporate yield to the estimated yield of a Treasury security of the same maturity.

To ensure that our results are not driven by a small number of extreme observations, we eliminated all observations with credit spreads below 5 basis points and greater than 3,500 basis points. In addition, we dropped from our sample very small corporate issues (par value of less than \$1 million) and all observations with a remaining term to maturity of less than 1 year or more than 30 years. These selection criteria yielded a sample of 5,982 individual securities for the period between January 1973 and September 2010. We matched these corporate securities with their issuer's quarterly income and balance sheet data from Compustat and daily data on equity valuations from CRSP, yielding a matched sample of 1,112 firms.

Table 1 contains summary statistics for the key characteristics of bonds in our sample. Note that a typical firm in our sample has only a few senior unsecured issues outstanding at any point in time—the median firm, for example, has two such issues trading in any given month. This distribution, however, exhibits a significant positive skew, as some firms can have many more issues trading in the secondary market at a point in time.

The distribution of the market values of these issues is similarly skewed, with the range running from \$1.2 million to more than \$5.6 billion. The maturity of these debt instruments is fairly long, with the average maturity at issue of 13 years; the average remaining term to maturity in our sample is 11.3 years. Because corporate bonds typically generate significant cash flows in the form of regular coupon payments, however, their average duration is considerably shorter.

An important characteristic of our sample is the fact that about two-thirds of the securities are callable—that is, the issuer has the right to “call” (i.e., redeem) the bond issue prior to its maturity under certain prespecified conditions. Moreover,

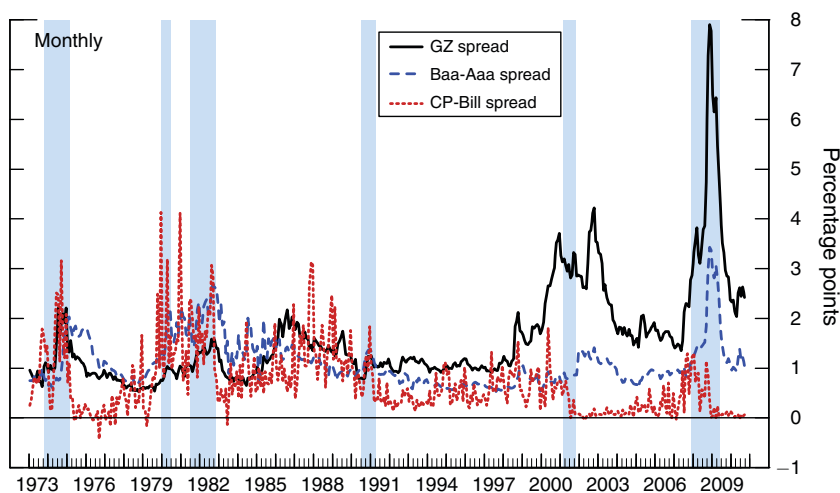


FIGURE 1. SELECTED CORPORATE CREDIT SPREADS

Notes: Sample period: 1973:1–2010:9. The figure depicts the following credit spreads: GZ spread = the average credit spread on senior unsecured bonds issued by nonfinancial firms in our sample (the solid line); Baa–Aaa = the spread between yields on Baa- and Aaa-rated long-term industrial corporate bonds (the dashed line); and CP–Bill = the spread between the yield on one-month A1/P1 nonfinancial commercial paper and the one-month Treasury yield (the dotted line). The shaded vertical bars represent the NBER-dated recessions.

the share of callable debt in the secondary market has varied substantially over the sample period, with almost all bonds being subject to a call provision until the late 1980s. Likely spurred by the decline in long-term nominal interest rates and the accompanied reduction in interest rate volatility, the share of callable debt fell to its historic low of about 25 percent by the mid-1990s. Over the past decade and a half, however, this trend has been almost completely reversed, as nonfinancial firms resumed issuing large amounts of callable senior unsecured debt.

In terms of default risk—at least as measured by the S&P credit ratings—our sample spans the entire spectrum of credit quality, from “single D” to “triple A.” At “BBB1,” however, the median observation is still solidly in the investment-grade category. An average bond has an expected return of 204 basis points above the comparable risk-free rate, while the sizable standard deviation of 281 basis points reflects the wide range of credit quality in our sample.

Using this micro-level dataset, we construct a simple credit-spread index that is representative of the entire maturity spectrum and the range of credit quality in the corporate cash market. Specifically, the GZ credit spread is calculated as

$$(1) \quad S_t^{GZ} = \frac{1}{N_t} \sum_i \sum_k S_{it}[k],$$

where N_t is the number of bond/firm observations in month t —that is, the GZ credit spread is simply an arithmetic average of the credit spreads on outstanding bonds in any given month. Figure 1 shows the GZ credit spread along with two widely used default-risk indicators that are also available over our sample period: the spread between yields on indexes of Baa- and Aaa-rated seasoned industrial corporate

bonds, and the yield spread between one-month A1/P1-rated nonfinancial commercial paper and the one-month Treasury yield (the paper–bill spread).³

All three credit spreads are clearly countercyclical, rising prior to and during economic downturns. Nonetheless, the pairwise correlations between the three series are fairly small and do not exhibit much of a systematic pattern. For example, the correlation between the paper–bill and the Baa–Aaa spreads is 0.21, whereas the paper–bill and the GZ spreads are slightly negatively correlated, with a correlation coefficient of -0.17 . Perhaps not too surprisingly, the highest correlation, 0.38, is between the two corporate bond credit spread indexes. Regarding their variability, the Baa–Aaa and the paper–bill spreads are the least volatile, with standard deviations of 50 and 67 basis points, respectively. Reflecting its broader coverage, both in terms of credit quality and maturity, the standard deviation of the GZ credit spread—at about 100 basis points—is considerably higher.

II. Credit Spreads and Economic Activity

To assess the predictive ability of credit spreads for economic activity (denoted by Y_t), we estimate the following univariate forecasting specification:

$$(2) \quad \nabla^h Y_{t+h} = \alpha + \sum_{i=1}^p \beta_i \nabla Y_{t-i} + \gamma_1 TS_t + \gamma_2 RFF_t + \gamma_3 CS_t + \epsilon_{t+h},$$

where $\nabla^h Y_{t+h} \equiv \frac{c}{h+1} \ln\left(\frac{Y_{t+h}}{Y_{t-1}}\right)$, $h \geq 0$ is the forecast horizon, and c is a scaling constant that depends on the frequency of the data (i.e., $c = 1,200$ for monthly data and $c = 400$ for quarterly data). In the forecasting regression (2), TS_t denotes the “term spread”—that is, the slope of the Treasury yield curve, defined as the difference between the three-month constant-maturity Treasury yield and the ten-year constant-maturity yield; RFF_t denotes the real federal funds rate; CS_t denotes a credit spread; and ϵ_{t+h} is the forecast error. Thus, our framework examines the marginal information content of credit spreads conditional on the slope of the yield curve and the real federal funds rate, two key indicators of the stance of monetary policy.⁴

The timing adopted by this specification allows for the possibility of “nowcasting” (i.e., $h = 0$), and it is intended to capture the fact that when forecasting an indicator of economic activity in period t , economists, because of reporting lags, typically do not observe the current value of the indicator, while the current financial asset prices are readily available. The forecasting regression (2) is estimated by ordinary least squares (OLS), with the lag length p of each specification determined by the Akaike Information Criterion (AIC). For the forecasting horizons $h \geq 1$, the MA(h) structure of the error term ϵ_{t+h} induced by overlapping observations is taken into account by computing standard errors according to Hodrick (1992).⁵

³ Other than than the GZ credit spread, all yields are taken from the “Selected Interest Rates” (H.15) statistical release published by the Federal Reserve Board.

⁴ The real federal funds rate in period t is defined as the average effective federal funds rate during period t less realized inflation, where realized inflation is given by the log-difference between the core PCE price index in period $t - 1$ and its lagged value a year earlier.

⁵ As shown by Ang and Bekaert (2007), the standard errors developed by Hodrick (1992) retain the correct size even in relatively small samples when testing the null of no predictability in the context of overlapping observations.

We first analyze the information content of the three credit spreads shown in Figure 1 for key monthly indicators of economic activity: the growth of private (nonfarm) payroll employment, the change in the civilian unemployment rate, and the growth of manufacturing industrial production. (In the case of the unemployment rate, the transformation ∇^h does not involve logs.) Using quarterly data, we also consider the predictive content of these default-risk indicators for the growth of real GDP, the broadest measure of economic activity.

A. Forecasting Results

The results in Table 2 detail the predictive power of various financial indicators for the three monthly measures of economic activity. We focus on the 3- and 12-month-ahead forecast horizons and report standardized estimates of the coefficients associated with the financial indicators, as well as the in-sample goodness of fit as measured by the adjusted R^2 .

The first column in each subpanel of the table contains the results from our baseline specification, which includes the term spread, the real federal funds rate, and p lags of ∇Y_{t-1} as predictors. In line with previous findings, the slope of the Treasury yield curve has significant predictive content for all three economic indicators at both forecast horizons, with a flat or inverted yield curve signaling a deterioration in labor market conditions and a deceleration in industrial output. The real federal funds rate has some additional predictive power for changes in the labor market conditions at the 12-month forecast horizon, but it has no explanatory power for the growth of industrial production at either horizon.

The remaining three columns contain the results from our baseline specification augmented with the three default-risk indicators. The paper-bill spread forecasts all three measures of economic activity, though the addition of this spread leads to only a small increase in the adjusted R^2 relative to the baseline specification. The forecasting ability of the Baa–Aaa credit spread appears to be equally modest; although the Baa–Aaa spread contains some marginal information for near-term economic developments, at the year-ahead horizon, this default-risk indicator has statistically significant, but economically negligible, explanatory power only for changes in the unemployment rate.

In contrast to the results obtained with the two standard default-risk indicators, the GZ credit spread is statistically a highly significant predictor of all three measures of economic activity at both the short- and longer-term horizons. Moreover, the magnitude of the estimated coefficients implies an economically significant negative relationship between credit spreads and future economic activity. For example, an increase of 100 basis points in the GZ credit spread in month t implies an almost 3.0 percentage point (annualized) drop in the growth rate of industrial output over the subsequent three months. Moreover, the inclusion of the GZ spread in the predictor set yields sizable improvements in the in-sample fit, ranging—at the 12-month horizon—from 12 percentage points in the case of industrial production to about 15 percentage points for the 2 labor market indicators.

Table 3 summarizes the predictive content of these financial indicators for the growth of real GDP. According to the entries in the table, the current stance of monetary policy has no predictive power for the next quarter's economic growth, although

TABLE 2—FINANCIAL INDICATORS AND MONTHLY MEASURES OF ECONOMIC ACTIVITY

Financial indicator	Forecast horizon: 3 months				Forecast horizon: 12 months			
<i>Panel A. Payroll employment</i>								
Term spread	-0.096 [2.12]	-0.102 [2.27]	-0.110 [2.44]	-0.110 [2.42]	-0.252 [4.94]	-0.255 [5.05]	-0.241 [4.93]	-0.277 [5.53]
Real FFR	-0.058 [1.18]	0.050 [0.78]	-0.038 [0.75]	-0.113 [2.23]	-0.116 [2.10]	-0.064 [0.96]	-0.129 [2.45]	-0.204 [3.84]
CP–bill spread	—	-0.165 [3.80]	—	—	—	-0.080 [2.29]	—	—
Baa–Aaa spread	—	—	-0.075 [2.05]	—	—	—	0.054 [1.15]	—
GZ spread	—	—	—	-0.322 [8.50]	—	—	—	-0.497 [13.4]
Adjusted R^2	0.622	0.639	0.625	0.685	0.422	0.424	0.422	0.579
<i>Panel B. Unemployment rate</i>								
Term spread	0.164 [7.71]	0.179 [8.42]	0.215 [10.2]	0.199 [9.37]	0.375 [46.7]	0.386 [48.4]	0.394 [50.6]	0.419 [51.8]
Real FFR	0.029 [1.24]	-0.152 [5.22]	-0.024 [1.06]	0.107 [4.50]	0.037 [4.60]	-0.089 [9.72]	0.019 [2.45]	0.131 [16.4]
CP–bill spread	—	0.268 [13.8]	—	—	—	0.191 [36.0]	—	—
Baa–Aaa spread	—	—	0.198 [10.4]	—	—	—	0.074 [11.2]	—
GZ spread	—	—	—	0.351 [19.5]	—	—	—	0.453 [83.0]
Adjusted R^2	0.335	0.378	0.362	0.425	0.270	0.292	0.273	0.417
<i>Panel C. Industrial production</i>								
Term spread	-0.182 [2.54]	-0.202 [2.83]	-0.239 [3.43]	-0.224 [3.15]	-0.358 [4.03]	-0.371 [4.91]	-0.357 [4.06]	-0.400 [4.59]
Real FFR	-0.035 [0.44]	0.183 [1.86]	0.016 [0.20]	-0.126 [1.62]	-0.094 [0.98]	0.052 [0.48]	-0.095 [1.02]	-0.175 [1.90]
CP–bill spread	—	-0.332 [4.75]	—	—	—	-0.226 [3.67]	—	—
Baa–Aaa spread	—	—	-0.211 [3.08]	—	—	—	0.004 [0.05]	—
GZ spread	—	—	—	-0.386 [5.28]	—	—	—	-0.412 [5.11]
Adjusted R^2	0.251	0.319	0.283	0.360	0.227	0.258	0.225	0.346

Notes: Sample period: 1973:1–2010:9. Dependent variable is $\nabla^h Y_{t+h}$, where Y_t denotes an indicator of economic activity in month t and h is the forecast horizon. In addition to the specified financial indicator in month t , each specification also includes a constant and p lags of ∇Y_{t-1} (not reported), where p is determined by the AIC. Entries in the table denote the standardized estimates of the OLS coefficients associated with each financial indicator; absolute asymptotic t -statistics reported in brackets are computed according to Hodrick (1992) (see text for details).

the term spread is economically and statistically a highly significant predictor of the year-ahead growth in real output. Both the paper–bill and the Baa–Aaa spreads contain some information about the near-term growth prospects, but the signaling ability of these two default-risk indicators vanishes at longer horizons. In contrast, the GZ credit spread is a highly significant predictor of real GDP growth at both the 1- and 4-quarter forecast horizons, with an increase of 100 basis points in the GZ credit spread in quarter t leading to a deceleration in real GDP of more than 1.25 percentage points over the subsequent 4 quarters.

TABLE 3—FINANCIAL INDICATORS AND REAL GDP

Financial indicator	Forecast horizon: 1 quarter				Forecast horizon: 4 quarters			
	Term spread	-0.198 [1.77]	-0.217 [1.92]	-0.250 [2.07]	-0.247 [2.26]	-0.398 [2.79]	-0.406 [2.81]	-0.413 [2.70]
Real FFR	-0.016 [0.12]	0.175 [1.12]	0.020 [0.15]	-0.123 [0.95]	-0.036 [0.24]	0.042 [0.22]	-0.026 [0.17]	-0.131 [0.87]
CP–bill spread	—	-0.254 [2.16]	—	—	—	-0.105 [0.82]	—	—
Baa–Aaa spread	—	—	-0.229 [1.95]	—	—	—	-0.066 [0.52]	—
GZ spread	—	—	—	-0.437 [4.96]	—	—	—	-0.482 [5.74]
Adjusted R^2	0.170	0.197	0.209	0.313	0.215	0.215	0.213	0.369

Notes: Sample period: 1973:I–2010:III. Dependent variable is $\nabla^h Y_{t+h}$, where Y_t denotes the real GDP in quarter t and h is the forecast horizon. In addition to the specified financial indicator in quarter t , each specification also includes a constant and p lags of ∇Y_{t-1} (not reported), where p is determined by the AIC. Entries in the table denote the standardized estimates of the OLS coefficients associated with each financial indicator; absolute asymptotic t -statistics reported in brackets are computed according to Hodrick (1992) (see text for details).

III. The Excess Bond Premium

In this section, we exploit the micro-level aspect of the data to decompose our high-information content credit spread index into two components: a component that captures the systematic movements in default risk of individual firms and a residual component—the excess bond premium—that, as we argue below, likely represents variation in the average price of bearing exposure to US corporate credit risk, above and beyond the compensation for expected defaults.

Our empirical methodology is related to the recent work of Berndt et al. (2008), in that the log of the credit spread on bond k (issued by firm i) at time t is assumed to be related linearly to a firm-specific measure of expected default DFT_{it} and a vector of bond-specific characteristics $\mathbf{Z}_{it}[k]$, according to

$$(3) \quad \ln S_{it}[k] = \beta DFT_{it} + \gamma' \mathbf{Z}_{it}[k] + \epsilon_{it}[k],$$

where the zero-mean disturbance $\epsilon_{it}[k]$ represents a “pricing error.”⁶ The credit-spread regression (3) is estimated by OLS, and the standard errors are double-clustered in the firm (i) and time (t) dimensions and thus are robust to both cross-sectional dependence and serial correlation; see, for example, Cameron, Gelbach, and Miller (2011).

Assuming normally distributed disturbances, the predicted level of the spread for bond k of firm i at time t is given by

$$\hat{S}_{it}[k] = \exp \left[\hat{\beta} DFT_{it} + \hat{\gamma}' \mathbf{Z}_{it}[k] + \frac{\hat{\sigma}^2}{2} \right],$$

⁶Taking logs of credit spreads provides a useful transformation to control for heteroskedasticity, given that the distribution of credit spreads is highly skewed.

where $(\hat{\beta}, \hat{\gamma}')$ denotes the OLS estimates of the corresponding parameters and $\hat{\sigma}^2$ is the estimated variance of the disturbance term $\epsilon_{it}[k]$. By averaging across bonds/firms at time t , we can define the predicted component of the GZ credit spread as

$$\hat{S}_t^{GZ} = \frac{1}{N_t} \sum_i \sum_k \hat{S}_{it}[k].$$

The excess bond premium in period t is then defined by the following linear decomposition:

$$EBP_t = S_t^{GZ} - \hat{S}_t^{GZ}.$$

Within this framework, we are interested in determining the extent to which the forecasting power of the GZ credit spread is due to the information content of the expected default component (\hat{S}^{GZ}) versus movements in the excess bond premium (EBP).

A. Measuring Default Risk

To measure a firm's probability of default at each point in time, we employ the "distance to default" (DD) framework developed in the seminal work of Merton (1974). The key insight of this contingent claims approach to corporate credit risk is that the equity of the firm can be viewed as a call option on the underlying value of the firm with a strike price equal to the face value of the firm's debt. Although neither the underlying value of the firm nor its volatility can be observed directly, they can, under the assumptions of the model, be inferred from the value of the firm's equity, the volatility of its equity, and the firm's observed capital structure.

While used widely by the financial industry, our choice of the Merton framework is also motivated by the work of Schaefer and Strebulaev (2008), who present compelling micro-level evidence showing that even the simplest structural default model—the DD model with nonstochastic interest rates—accounts well for the default-risk component of corporate bond prices. In particular, such models generate sensitivities (i.e., hedge ratios) of corporate bond returns to the issuing firm's equity and riskless bond returns that are remarkably consistent with those observed in the actual data.

The first assumption underlying the DD framework is that the total value of the firm V follows a geometric Brownian motion:

$$(4) \quad dV = \mu_V V dt + \sigma_V V dW,$$

where μ_V denotes the expected continuously compounded return on V ; σ_V is the volatility of firm value; and dW is an increment of the standard Weiner process. The second assumption pertains to the firm's capital structure. In particular, it is assumed that the firm has just issued a single discount bond in the amount D that will mature in T periods.

Together, these two assumption imply that the value of the firm's equity E can be viewed as a call option on the underlying value of the firm V , with a strike price

equal to the face value of the firm's debt D and a time to maturity of T . According to the Black-Scholes-Merton option-pricing framework, the value of the firm's equity then satisfies

$$(5) \quad E = V\Phi(\delta_1) - e^{-rT}D\Phi(\delta_2),$$

where r denotes the instantaneous risk-free interest rate, $\Phi(\cdot)$ is the cumulative standard normal distribution function, and

$$\delta_1 = \frac{\ln(V/D) + (r + 0.5\sigma_V^2)T}{\sigma_V\sqrt{T}} \quad \text{and} \quad \delta_2 = \delta_1 - \sigma_V\sqrt{T}.$$

According to equation (5), the value of the firm's equity depends on the total value of the firm and time, a relationship that also underpins the link between the volatility of the firm's value σ_V and the volatility of its equity σ_E . In particular, it follows from Ito's lemma that

$$(6) \quad \sigma_E = \left[\frac{V}{E} \right] \frac{\partial E}{\partial V} \sigma_V.$$

Under the Black-Scholes-Merton option-pricing framework $\frac{\partial E}{\partial V} = \Phi(\delta_1)$, the relationship between the volatility of the firm's value and the volatility of its equity is given by

$$(7) \quad \sigma_E = \left[\frac{V}{E} \right] \Phi(\delta_1) \sigma_V.$$

From an operational standpoint, the most critical inputs to the Merton DD model are clearly the market value of equity E , the face value of debt D , and the volatility of equity σ_E . Assuming a forecasting horizon of 1 year (i.e., $T = 1$), we implement the model in 2 steps: First, we estimate σ_E from historical daily stock returns using a 250-day moving window. Second, we assume that the face value of the firm's debt D is equal to the sum of the firm's current liabilities and one-half of its long-term liabilities.⁷ Using the observed values of E , D , σ_E , and r (i.e., the daily one-year constant-maturity Treasury yield), equations (5) and (7) can be solved for V and σ_V using standard numerical techniques.

As noted by Vassalou and Xing (2004), for example, the excessive volatility of market leverage (V/E) in equation (7) causes large swings in the estimated volatility of the firm's value σ_V , which are difficult to reconcile with the observed frequency of defaults and movements in financial asset prices. To resolve this problem, we implement an iterative procedure proposed by Bharath and Shumway (2008). The procedure involves the following steps: First, we initialize the procedure by

⁷This assumption for the "default point" is also used by Moody's/KMV in the construction of their expected default frequencies (EDFs), which are based on the DD framework. The assumption captures the notion that short-term debt requires a repayment of the principal relatively soon, whereas long-term debt requires the firm to meet only the coupon payments. Both current and long-term liabilities are taken from quarterly Compustat files and interpolated to daily frequency using a step function.

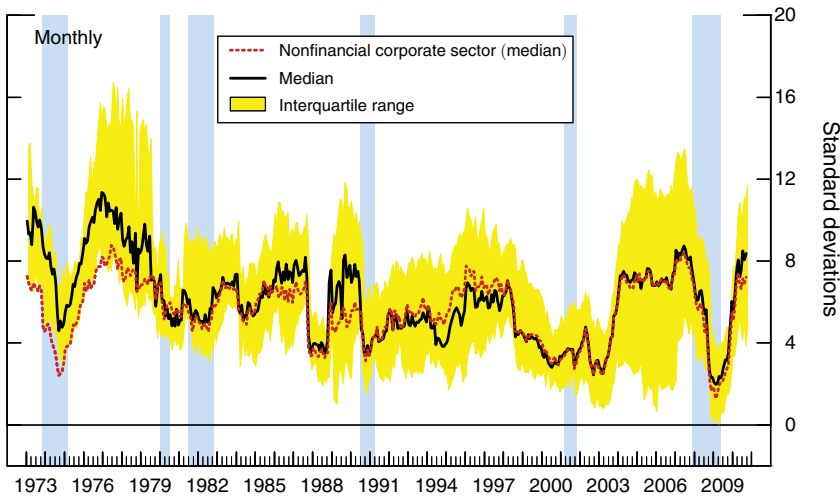


FIGURE 2. DISTANCE TO DEFAULT

Notes: Sample period: 1973:1–2010:9. The solid line depicts the (weighted) median DD of the firms in our sample, and the shaded band depicts the corresponding (weighted) interquartile range. The dotted line depicts the (weighted) median DD in the US nonfinancial corporate sector; all percentiles are weighted by the firm's outstanding liabilities. The shaded vertical bars represent the NBER-dated recessions.

letting $\sigma_V = \sigma_E[D/(E + D)]$. We then use this value of σ_V in equation (5) to infer the market value of the firm's assets V for every day of the 250-day moving window. In the second step, we calculate the implied daily log-return on assets (i.e., $\Delta \ln V$) and use the resulting series to generate new estimates of σ_V and μ_V . We then iterate on σ_V until convergence.

The resulting solutions of the Merton DD model can be used to calculate the firm-specific DD over the one-year horizon as

$$(8) \quad DD = \frac{\ln(V/D) + (\mu_V - 0.5\sigma_V^2)}{\sigma_V}.$$

In this context, default occurs when the ratio of the value of assets to debt in equation (8) falls below one (or its log is negative); in effect, DD measures the number of standard deviations the log of this ratio must deviate from its mean for default to occur. The implied probability of default is given by $\Phi(-DD)$, which, under the assumptions of the model, should be a sufficient statistic for predicting defaults.

Using this methodology, we compute the year-ahead DD for all US nonfinancial corporations covered by the S&P's Compustat and CRSP over our sample period. Figure 2 plots the cross-sectional median and the interquartile range of the DD for the 1,112 bond issuers in our sample. As a point of comparison, the figure also depicts the cross-sectional median of the DD for the entire Compustat-CRSP matched sample (14,458 firms) of nonfinancial firms.⁸ The median DD for both sets

⁸We eliminated from our sample all observations with the DD of less than -2 or more than 20 , cutoffs corresponding roughly to the first and 99th percentiles of the DD distribution, respectively.

of firms is strongly procyclical, implying that equity investors generally expect an increase in defaults during economic downturns. Indeed, during the height of the recent financial crisis in the autumn of 2008, both measures fell to record lows, a drop consistent with the jump in the GZ credit spread shown in Figure 1.

B. Credit Spreads and Default Risk

With our firm-specific measure of default risk in hand, we now turn to the estimation of the credit-spread model given in equation (3). In our baseline specification, we regress $\ln S_{it}[k]$, the logarithm of the credit spread on bond k (issued by firm i) in month t , on the distance to default DD_{it} , while also controlling for bond-specific characteristics that could influence bond yields through either term or liquidity premiums. These predetermined characteristics, denoted by the vector $\mathbf{Z}_{it}[k]$, include the bond's duration ($DUR_{it}[k]$), the amount outstanding ($PAR_{it}[k]$), the (fixed) coupon rate ($CPN_{it}[k]$), the age of the issue ($AGE_{it}[k]$), and an indicator variable that equals one if the bond is callable and zero otherwise ($CALL_{it}[k]$).

The regression also includes three-digit NAICS industry fixed effects to control for any systematic (time-invariant) differences in expected recovery rates across industries. Lastly, the specification includes S&P credit rating fixed effects, which capture the “soft information” regarding the firm's financial health, information that is complementary to our option-theoretic measures of default risk, according to Löffler (2004, 2007).⁹

As shown in the left panel of Table 4, our market-based measure of default risk is statistically a highly significant predictor of the log credit spreads. In economic terms, the estimated coefficient on the distance to default implies that a decrease of one standard deviation in the year-ahead DD leads to a widening of credit spreads of about 15 basis points. As evidenced by the adjusted R^2 , the baseline credit-spread model explains a considerable portion of the variation in the log credit spreads.

The DD should summarize all available information regarding the risk of default, according to the Merton model. Consequently, movements in risk-free interest rates should affect credit spreads only insofar as they change expected future cash flows and, as a result, the distance to default. As shown by Duffee (1998), however, if the firm's outstanding bonds are callable, then movements in risk-free rates—by changing the value of the embedded call option—will have an independent effect on bond prices, complicating the interpretation of the behavior of credit spreads. In addition, callable bonds are likely to be more sensitive to uncertainty regarding the future course of interest rates. On the other hand, to the extent that callable bonds are, in effect, of shorter duration, they may be less sensitive to changes in default risk.

⁹We conducted sensitivity analysis by adding quadratic (and higher-order) terms of the distance to default to our baseline specification in order to allow for a nonlinear effect of leverage on credit spreads; see, for example, Levin, Natalucci, and Zakrajšek (2004). Note also that according to equation (8), the distance to default consists of the log of leverage, the expected return on assets, and the volatility of asset returns. When estimating the excess bond premium, these three terms are constrained to enter the credit-spread regression through their effect on the distance to default. To the extent that the distance to default is not a sufficient statistic for default risk, these terms may have independent effects on the credit spreads that should be accounted for in the estimation. As an additional robustness check, we estimated a specification in which the three components of the distance to default—and their interactions with the call-option indicator—were allowed to separately affect the credit spreads. These modifications, however, had virtually no effect on any of our results.

TABLE 4—CREDIT SPREADS AND THE DISTANCE TO DEFAULT

Explanatory variable	Est.	Standard error	Est.	Standard error
$-DD_{it}$	0.075	0.005	0.093	0.005
$\ln(DUR_{it}[k])$	0.106	0.018	0.201	0.019
$\ln(PAR_{it}[k])$	0.171	0.018	0.121	0.022
$\ln(CPN_{it}[k])$	0.439	0.074	0.031	0.062
$\ln(AGE_{it}[k])$	0.047	0.008	0.135	0.010
$CALL_{it}[k]$	0.262	0.029	-0.427	0.210
$-DD_{it} \times CALL_{it}[k]$	—	—	-0.030	0.004
$\ln(DUR_{it}[k]) \times CALL_{it}[k]$	—	—	-0.120	0.023
$\ln(PAR_{it}[k]) \times CALL_{it}[k]$	—	—	-0.122	0.024
$\ln(CPN_{it}[k]) \times CALL_{it}[k]$	—	—	0.915	0.078
$\ln(AGE_{it}[k]) \times CALL_{it}[k]$	—	—	-0.132	0.013
$LEV_t \times CALL_{it}[k]$	—	—	-0.385	0.027
$SLP_t \times CALL_{it}[k]$	—	—	-0.088	0.017
$CRV_t \times CALL_{it}[k]$	—	—	-0.041	0.019
$VOL_t \times CALL_{it}[k]$	—	—	0.134	0.021
Adjusted R^2		0.649		0.700
Industry effects ^a		0.000		0.000
Credit rating effects ^b		0.000		0.000

Notes: Sample period: 1973:1–2010:9. Obs. = 346,126; Number of bonds/firms = 5,982/1,112. Dependent variable is $\ln(S_{it}[k])$, the log of the credit spread on bond k (issued by firm i) in month t . The Treasury term structure is represented by the following three factors: LEV_t = level; SLP_t = slope; and CRV_t = curvature; VOL_t = (annualized) realized monthly volatility of the daily ten-year Treasury yield. Asymptotic standard errors are clustered in both the firm (i) and time (t) dimensions, according to Cameron, Gelbach, and Miller (2011).

^a p -value of the exclusion test of industry fixed effects.

^b p -value of the exclusion test of credit rating fixed effects.

One possible way to deal with this issue would be to confine the analysis to a subsample of noncallable bonds. As reported in Table 1, however, callable bonds account, on average, for two-thirds of the senior unsecured corporate debt traded in the secondary market. Moreover, given the variation in the share of callable debt over time, limiting the sample to noncallable bonds would severely limit the time span of our data, making it impossible to shed much light on the recent financial crisis.

As an alternative, we control directly for the effects of the Treasury term structure and interest rate volatility on credit spreads of callable bonds when estimating the excess bond premium. In addition to interacting the distance to default and the vector of bond characteristics $\mathbf{Z}_{it}[k]$ in equation (3) with the $CALL_{it}[k]$ indicator, the credit spreads of callable bonds are also allowed to depend on the level, slope, and curvature of the Treasury yield curve, the three factors that summarize the vast majority of the information in the term structure, according to Litterman and Scheinkman (1991); the credit spreads of callable bonds can also be affected by the realized monthly volatility of the daily ten-year Treasury yield, a proxy for interest rate uncertainty.¹⁰

¹⁰The level, slope, and curvature factors correspond, respectively, to the first three principal components of nominal Treasury yields at 3-month, 6-month, and 1-, 2-, 3-, 5-, 7-, 10-, 15-, and 30-year maturities. All yield series are monthly (at month-end) and with the exception of the three- and six-month bill rates are derived from the smoothed Treasury yield curve estimated by Gürkaynak, Sack, and Wright (2007).

TABLE 5—SELECTED MARGINAL EFFECTS BY TYPE OF BOND

Explanatory variable	Noncallable		Callable		Mean ^a	STD ^b
	Est.	Standard error	Est.	Standard error		
Distance to default: $-DD_{it}$	0.190	0.010	0.129	0.008	6.610	3.946
Term structure: LEV_t	—	—	-0.783	0.055	0.000	1.000
Term structure: SLP_t	—	—	-0.179	0.034	0.000	1.000
Term structure: CRV_t	—	—	-0.082	0.038	0.000	1.000
Term structure: VOL_t (%)	—	—	0.273	0.043	1.862	1.239

Notes: The table contains the estimates of the marginal effect of a one-unit change in the specified variable on the level of credit spreads (in percentage points) for noncallable and callable bonds based on the parameter estimates reported in Table 4. All marginal effects are evaluated at sample means; by construction, the level, slope, and curvature factors are standardized to have mean equal to zero and standard deviation equal to one. Asymptotic standard errors are computed according to the delta method.

^a Sample mean of the specified variable.

^b Sample standard deviation of the specified variable.

The results of this exercise are reported in the right panel of Table 4. As predicted by the theory, an increase in the general level of interest rates and the steepening of the Treasury term structure—the effects captured by the level and slope factors, respectively—lead to a narrowing of credit spreads of callable bonds. In contrast, an increase in the realized volatility of longer-term Treasury yields boosts the spreads of callable bonds. Importantly, the inclusion of the term structure and volatility factors noticeably improves the fit of the credit-spread regression.

In Table 5, we translate the coefficients from the estimated log-spread pricing equation into the impact of variation in default risk, the shape of the term structure, and interest rate volatility on the level of credit spreads. In line with the theoretical predictions, the effect of default risk on credit spreads of callable bonds is significantly attenuated by the call-option mechanism, with a one standard deviation decline in the distance to default implying an increase of 29 basis points in the spreads of noncallable bonds, compared with a 13-basis-point rise in the spreads of their callable counterparts.

Consistent with the results of Duffee (1998), our estimates also imply that the shape of the Treasury term structure and interest rate volatility have economically significant effects on credit spreads of callable bonds. For example, a one-standard-deviation increase in the level factor implies a reduction in the credit spreads on callable bonds of almost 80 basis points, while a one-standard-deviation increase in the slope factor lowers credit spreads on such bonds 18 basis points. An increase in the volatility of long-term interest rates—by boosting the value of embedded call options—implies a widening of callable credit spreads of 27 basis points.

Figure 3 shows the GZ credit spread along with the fitted values from two specifications: one that includes the effects of the term structure terms on credit spreads of callable bonds and one that does not. Over most of our sample period, the option adjustment for callable bonds has had relatively little effect. One exception is the 1979–82 period of nonborrowed reserves targeting, a period characterized by a substantial volatility in nominal interest rates. Given that most of the bonds in our sample during that period were callable, increased interest rate volatility implies a higher fitted average spread, relative to the fitted value that does not control for interest rate volatility; in addition, the excessive volatility of credit spreads during this period implies more volatile fitted values.

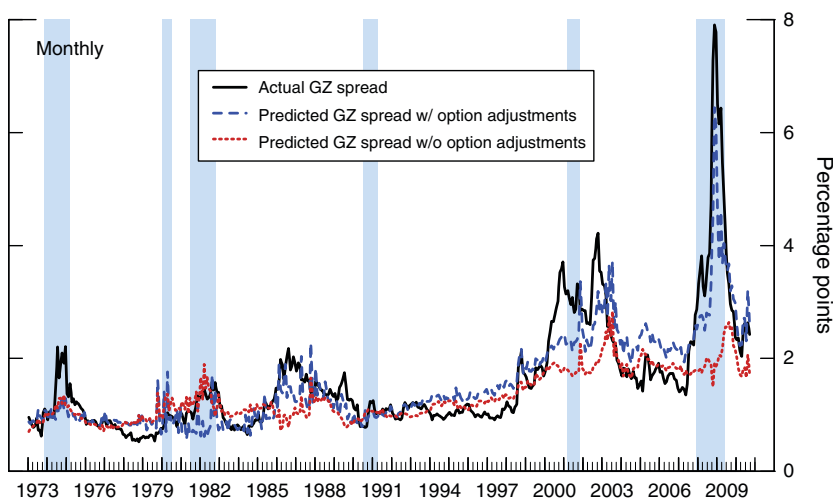


FIGURE 3. ACTUAL AND PREDICTED CREDIT SPREADS

Notes: Sample period: 1973:1–2010:9. The solid line depicts the actual GZ credit spread. The dashed line depicts the predicted GZ credit spread based on the specification that includes the term structure option-adjustment terms; the dotted line depicts the predicted GZ credit spread based on the specification that excludes the term structure option-adjustment terms (see text for details). The shaded vertical bars represent the NBER-dated recessions.

The option adjustment also had a significant effect during the recent financial crisis, reflecting the fact that the general level of interest rates fell to historically low levels. Because a low level of interest rates implies higher predicted values for the credit spreads of callable bonds, our option-adjustment procedure accounts for about 200 basis points of the total increase in the GZ credit spread during the height of the financial crisis in the autumn of 2008. Overall, the fitted values from this specification capture a substantial fraction of cyclical fluctuations in the GZ credit spread.

Figure 4 shows the estimated excess bond premium—that is, the difference between the GZ credit spread and the fitted value from the second specification in Table 4. With the exception of the 1990–1991 recession, the premium increased significantly prior to or during all cyclical downturns. The excess bond premium fell to a historically low level in the latter part of 2003 and remained low during the following several years, the period that, at least in retrospect, has been characterized by lax credit standards, excessive credit growth, and unsustainable asset price appreciation.

The intensification of credit concerns in US and foreign financial markets during the summer of 2007 precipitated a sharp increase in the excess bond premium, which continued to increase throughout the subsequent financial crisis, reaching a record high of 275 basis points in October 2008. Although conditions in financial markets improved somewhat over the remainder of 2008, investors' concern in early 2009 about the viability of major financial institutions led to another surge in the excess bond premium. Since then, this gauge of financial disruptions has reversed all of its run-up, a pattern consistent with the improved economic outlook and the easing of strains in financial markets.

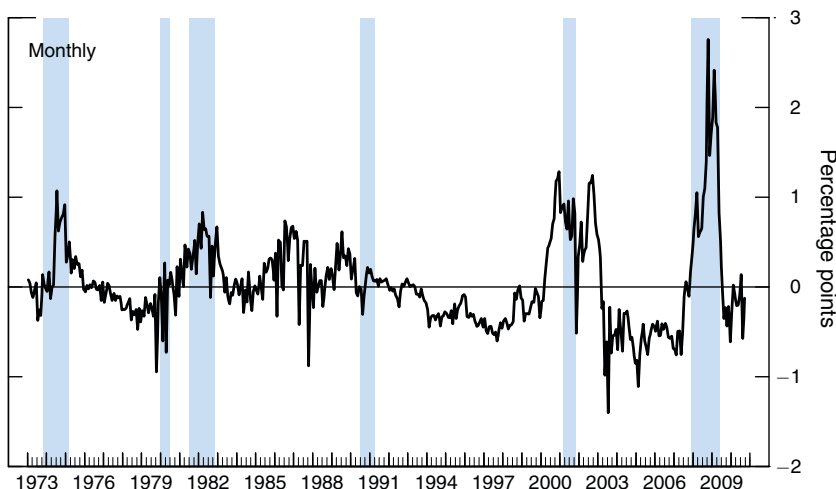


FIGURE 4. THE EXCESS BOND PREMIUM

Notes: Sample period: 1973:1–2010:9. The figure depicts the estimated (option-adjusted) excess bond premium. The shaded vertical bars represent the NBER-dated recessions.

IV. The Excess Bond Premium and Economic Activity

Our decomposition of the GZ credit spread implies that an important component of the variation in corporate credit spreads is due to fluctuations in the excess bond premium. We now examine whether movements in the excess bond premium provide independent information about future economic activity. First, we analyze the extent to which the forecasting power of the GZ credit spread documented in Section III is attributable to its predicted component or the excess bond premium. We then add the excess bond premium to an otherwise standard macroeconomic VAR and examine the implications of innovations to the excess bond premium for the real economy and asset prices.

A. Forecasting Results

Table 6 reports the results for the monthly indicators of economic activity, based on the specification in which the two components of the GZ credit spread— \hat{S}^{GZ} and EBP —are allowed to enter the forecasting regression (2) separately. According to our estimates, both the excess bond premium and the predicted GZ credit spread contain significant independent explanatory power for all 3 economic indicators, at both the 3- and 12-month forecast horizons. The (absolute) magnitude of the estimated coefficients on the excess bond premium, however, tends to be significantly larger than that of the coefficients associated with the predicted GZ spread, a finding indicating that the information content of credit spreads for economic activity largely reflects fluctuations in the nondefault component of credit spreads as opposed to movements in expected defaults.

In Table 7, we repeat this forecasting exercise for the growth rate of real GDP and its main components. To conserve space, we report the results for the four-quarter

TABLE 6—THE EXCESS BOND PREMIUM AND MONTHLY MEASURES OF ECONOMIC ACTIVITY

Financial indicator	Forecast horizon: 3 months			Forecast horizon: 12 months		
	EMP	UER	IPM	EMP	UER	IPM
Term spread	-0.122 [2.67]	0.221 [10.3]	-0.256 [3.60]	-0.291 [5.78]	0.440 [55.4]	-0.435 [4.99]
Real FFR	-0.044 [0.87]	0.007 [0.30]	-0.018 [0.23]	-0.112 [2.06]	0.022 [2.70]	-0.079 [0.81]
Predicted GZ spread ^a	-0.202 [5.65]	0.134 [8.41]	-0.186 [3.62]	-0.355 [9.63]	0.213 [38.5]	-0.283 [4.18]
Excess bond premium	-0.259 [8.52]	0.331 [20.9]	-0.386 [5.87]	-0.369 [14.5]	0.414 [91.6]	-0.388 [5.42]
Adjusted R^2	0.687	0.430	0.381	0.588	0.433	0.384

Notes: Sample period: 1973:1–2010:9. Dependent variable is $\nabla^h Y_{t+h}$, where Y_t denotes an indicator of economic activity in month t and h is the forecast horizon: EMP = private nonfarm payroll employment; UER = civilian unemployment rate; and IPM = index of manufacturing industrial production. In addition to the specified financial indicators in month t , each specification also includes a constant and p lags of ∇Y_{t-1} (not reported), where p is determined by the AIC. Entries in the table denote the standardized estimates of the OLS coefficients associated with each financial indicator; absolute asymptotic t -statistics reported in brackets are computed according to Hodrick (1992) (see text for details).

^aExcludes the effect of option adjustment on callable bonds.

forecast horizon only. We do, however, perform an important robustness check by performing the analysis for the 1985–2010 subsample, a period characterized by a stable monetary policy regime and by significant deregulation of financial markets.¹¹

As indicated in the first column of panel A of Table 7, the excess bond premium is economically and statistically a highly significant predictor of output growth at the year-ahead forecast horizon over the full sample period. The coefficient estimate implies that an increase in the excess bond premium of 100 basis points in quarter t leads to a drop in real GDP growth of more than 1.5 percentage points over the subsequent 4 quarters. Consistent with our previous findings, the impact on economic growth of a similarly sized move in the predicted component of the GZ credit spread is considerably smaller—a 100-basis-point increase implies a deceleration in output of only 0.5 percentage point.

The remaining columns in the top panel focus on the main categories of personal consumption expenditures and private investment. The excess bond premium has substantial predictive content for the growth of consumption spending on nondurables and services, the major components of business fixed investment, as well as for inventory accumulation, an especially volatile component of aggregate demand. With the exception of high-tech investment, the coefficients on the predicted component of the GZ spread are considerably smaller (in absolute value) than the respective coefficients on the excess bond premium, again indicating that movements in the excess bond premium have, in economic terms, a greater impact on aggregate economic activity. In fact, for the most cyclically volatile series such as inventory

¹¹Formal statistical tests of the stability of the forecasting regression function do indicate a possible structural break in the coefficients associated with financial indicators—most notably for the coefficient on the real federal funds rate. Given the well-documented change in the monetary policy operating procedures that took place during the late 1970s and the early 1980s, splitting the sample in 1985 thus provides a natural point to examine the robustness of our results across different sample periods.

TABLE 7—THE EXCESS BOND PREMIUM, REAL GDP, AND ITS MAIN COMPONENTS

Financial indicator	GDP	C-NDS	C-D	I-RES	I-ES	I-HT	I-NRS	INV
<i>Panel A. 1973:I–2010:III (Forecast horizon: Four quarters)</i>								
Term spread	−0.478 [3.33]	−0.452 [3.89]	−0.551 [2.55]	−0.564 [5.23]	−0.398 [3.16]	−0.098 [0.83]	0.317 [2.73]	−0.123 [1.43]
Real FFR	−0.036 [0.24]	0.106 [0.99]	0.106 [0.58]	−0.003 [0.03]	−0.086 [0.82]	−0.092 [0.67]	−0.111 [0.87]	0.014 [0.15]
Predicted GZ spread ^a	−0.258 [2.56]	−0.209 [2.39]	0.014 [0.11]	−0.159 [2.10]	−0.221 [2.48]	−0.426 [4.43]	−0.186 [2.07]	−0.287 [4.11]
Excess bond premium	−0.364 [5.36]	−0.260 [4.36]	−0.127 [1.00]	−0.018 [0.29]	−0.558 [5.87]	−0.374 [4.42]	−0.587 [5.77]	−0.656 [9.39]
Adjusted R^2	0.365	0.349	0.224	0.419	0.481	0.432	0.557	0.580
<i>Panel B. 1985:I–2010:III (Forecast horizon: Four quarters)</i>								
Term spread	−0.509 [4.09]	−0.362 [4.07]	−0.456 [2.00]	−0.596 [6.52]	−0.340 [2.75]	−0.071 [0.60]	0.392 [2.46]	−0.321 [4.08]
Real FFR	0.424 [2.67]	0.181 [1.53]	0.395 [1.49]	0.331 [2.96]	0.032 [0.18]	−0.130 [0.90]	0.000 [0.00]	0.301 [2.47]
Predicted GZ spread ^a	−0.023 [0.20]	−0.093 [0.82]	0.194 [0.99]	0.045 [0.47]	−0.088 [0.63]	−0.294 [2.11]	−0.061 [0.49]	−0.047 [0.46]
Excess bond premium	−0.501 [6.80]	−0.362 [5.02]	−0.260 [1.51]	−0.035 [0.50]	−0.650 [5.41]	−0.382 [3.39]	−0.613 [4.76]	−0.701 [8.20]
Adjusted R^2	0.357	0.508	0.101	0.484	0.448	0.405	0.624	0.635

Notes: Dependent variable is $\nabla^4 Y_{t+4}$, where Y_t denotes real GDP or one of its components in quarter t ; C-NDS = PCE on nondurable goods & services; C-D = PCE on durable goods; I-RES = residential investment; I-ES = business fixed investment in E&S (excluding high tech); I-HT = business fixed investment in high-tech equipment; I-NRS = business fixed investment in structures; INV = business inventories. In addition to the specified financial indicators in quarter t , each specification also includes a constant and p lags of ∇Y_{t-1} (not reported), where p is determined by the AIC. Entries in the table denote the standardized estimates of the OLS coefficients associated with each financial indicator; absolute asymptotic t -statistics reported in brackets are computed according to Hodrick (1992) (see text for details).

^aExcludes the effect of option adjustment on callable bonds.

investment and spending on equipment and software (E&S) and nonresidential structures, the economic impact of the excess bond premium is more than twice as large as that of the predicted component of the GZ credit spread.

As shown in Panel B, the predictive content of the excess bond premium for economic activity over the 1985–2010 period is, if anything, greater than that obtained for the full sample period. The forecasting ability of the excess bond premium over the latter subsample is especially striking in the case of real GDP growth. According to our estimates, the predicted component of the GZ credit spread has no forecasting power for the growth of real GDP since the mid-1980s, while the excess bond premium continues to provide economically and statistically highly significant signals regarding economic growth prospects. In general, the coefficients on the excess bond premium estimated over this subperiod are noticeably higher (in absolute value) than those reported in the top panel. The estimates based on the 1985–2010 period imply that a 100-basis-point increase in the excess bond premium in quarter t lowers output about 2 percentage points over the next 4 quarters.

In summary, the above analysis indicates that the excess bond premium is a robust predictor of economic activity. This finding holds true across a variety of economic indicators, short- and longer-term forecast horizons, and sample periods. Furthermore, our forecasting results imply that since the mid-1980s, most of the

predictive content of the GZ credit spread for economic activity can be attributed to variation in the excess bond premium rather than to variation in default risk, as measured by its predicted component.

B. Macroeconomic Implications

In this section, we examine the macroeconomic consequences of shocks to the excess bond premium. We do so by adding the excess bond premium to a standard VAR that includes the following endogenous variables: (i) log-difference of real personal consumption expenditures (PCE); (ii) log-difference of real business fixed investment (BFI); (iii) log-difference of real GDP; (iv) inflation as measured by the log-difference of the GDP price deflator; (v) the quarterly average of the excess bond premium; (vi) the quarterly (value-weighted) excess stock market return from CRSP; (vii) the ten-year (nominal) Treasury yield; and (viii) the effective (nominal) federal funds rate. The identifying assumption implied by this recursive ordering is that shocks to the excess bond premium affect economic activity and inflation with a lag, while the risk-free rates and stock prices can react contemporaneously to such a financial disturbance; the VAR is estimated over the full sample period, using two lags of each endogenous variable.

Figure 5 depicts the impulse response functions of the endogenous variables to an orthogonalized shock to the excess bond premium. An unanticipated increase of one standard deviation in the excess bond premium—about 20 basis points—causes a significant reduction in real economic activity, with consumption, investment, and output all falling over the next several quarters. The macroeconomic consequences of this adverse financial shock are substantial; the level of real GDP bottoms out about 0.5 percentage point below trend five quarters after the shock, while the drop in investment is much more severe and persistent.

The resulting economic slack leads to a substantial disinflation over time. In response to these adverse economic developments, monetary policy is eased significantly, as evidenced by the decline in the federal funds rate that commences about one quarter after the initial impact of the shock. Despite the reduction in the overnight policy rate and the associated decline in longer-term yields, the stock market experiences a significant drop, with cumulative decline of about 7.0 percentage points relative to trend growth.¹²

¹²The quantitative and qualitative nature of these results is robust to an alternative identification scheme, in which the excess bond premium is ordered last in the “fast moving” financial block of the VAR. As an additional robustness check, we also estimated a similar VAR using monthly data, which renders our identifying assumptions—especially in the case of slow moving economic variables—considerably more realistic and provides an even more plausible identification of credit supply shocks. The impulse responses of the various measures of economic activity and inflation from the monthly VAR were very similar to those in shown in Figure 5. Moreover, when ordered first among the financial variables, the contemporaneous impact of the orthogonalized shock to the excess bond premium on financial variables is quite small in economic terms and in many cases statistically indistinguishable from zero. All told, our VAR results are robust to alternative identification assumptions as well as monthly versus quarterly data. In addition, restricting our sample to the post-1985 period indicates that the macroeconomic effects of a shock to the excess bond premium are, if anything, larger than those estimated over the full sample period. This finding is consistent with the results reported in panel B of Table 7, which show that over the 1985–2010 period, movements in the excess bond premium appear to have been especially informative about the subsequent fluctuations in the cyclically sensitive components of aggregate demand.

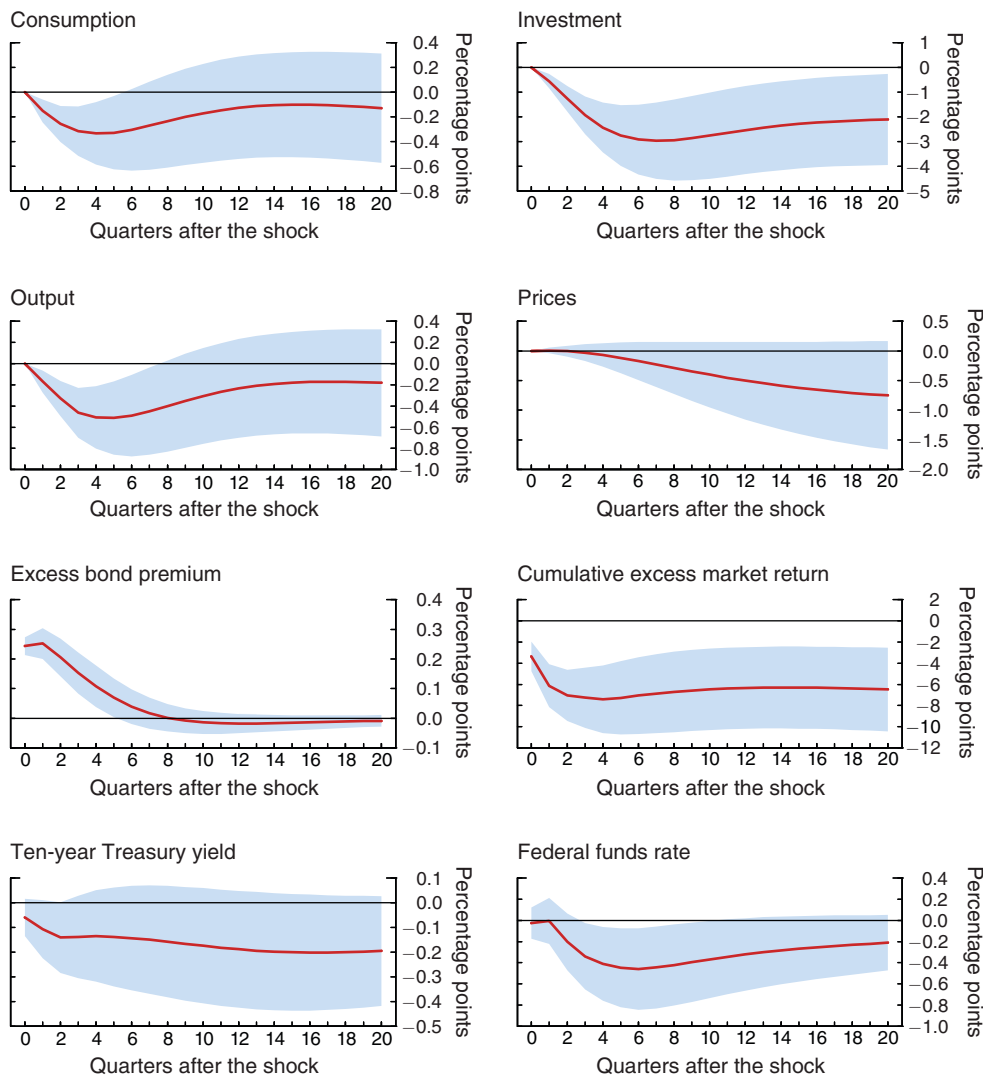


FIGURE 5. MACROECONOMIC IMPLICATIONS OF A FINANCIAL SHOCK

Notes: The figure depicts the impulse responses to a one-standard-deviation orthogonalized shock to the excess bond premium (see text for details). The responses of consumption, investment, and output growth and that of the excess market return have been accumulated. Shaded bands denote 95-percent confidence intervals based on 2,000 bootstrap replications.

Figure 6 shows the amount of variation in the endogenous variables explained by the orthogonalized shocks to the excess bond premium. These innovations account for more than 10 percent of the variation in output and 25 percent of the variation in business fixed investment at business cycle frequencies, proportions that exceed the amount of variation typically explained by monetary policy shocks. In addition, shocks to the excess bond premium explain a significant portion of the variation in broad equity valuations.

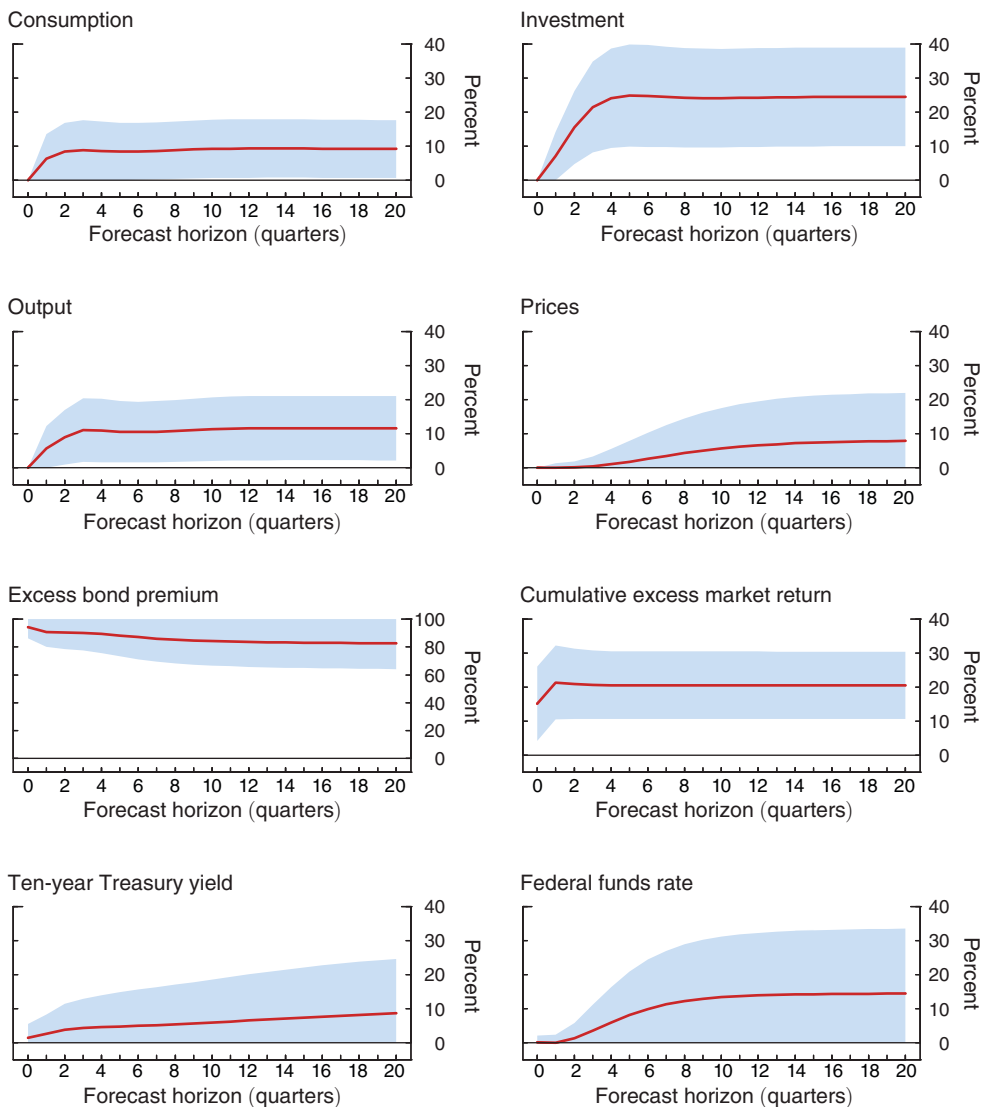


FIGURE 6. FORECAST ERROR VARIANCE DECOMPOSITION OF A FINANCIAL SHOCK

Notes: The figure depicts the forecast error variance decomposition from a one-standard-deviation orthogonalized shock to the excess bond premium (see text for details). The forecast error variance decomposition of consumption, investment, and output growth and that of the excess market return is based on the level of the variables. Shaded bands denote 95-percent confidence intervals based on 2,000 bootstrap replications.

C. Interpretation

The macroeconomic dynamics reported above are consistent with the notion that the excess bond premium provides a timely and useful gauge of credit supply conditions. A reduction in the supply of credit—an increase in the excess bond premium—causes a drop in asset prices and a contraction in economic activity through the financial accelerator mechanisms emphasized by Bernanke and Gertler (1989); Kiyotaki and Moore (1997); Bernanke, Gertler, and Gilchrist (1999); and

Hall (2011). Our findings also provide empirical support for the recent work of Gertler and Kiyotaki (2010); Gertler and Karadi (2011); and Gertler, Kiyotaki, and Queralto (forthcoming), who introduce macroeconomic models in which shocks to the value of assets held by financial intermediaries—by reducing the supply of credit—have independent effects on the real economy.

Given the inherent asymmetric feature of debt contracts, our results could also reflect the fact that prices of corporate bonds—compared with equity prices—are better able to capture the downside risks to economic growth. Thus, fluctuations in the excess bond premium may be due in part to a small but time-varying risk of economic “disasters.” As shown recently by Gourio (2010), a small increase in the probability of such an extreme event can cause a collapse in investment in a canonical real business cycle framework, *vis-à-vis* a sharp increase in the risk premium that significantly boosts the cost of capital.

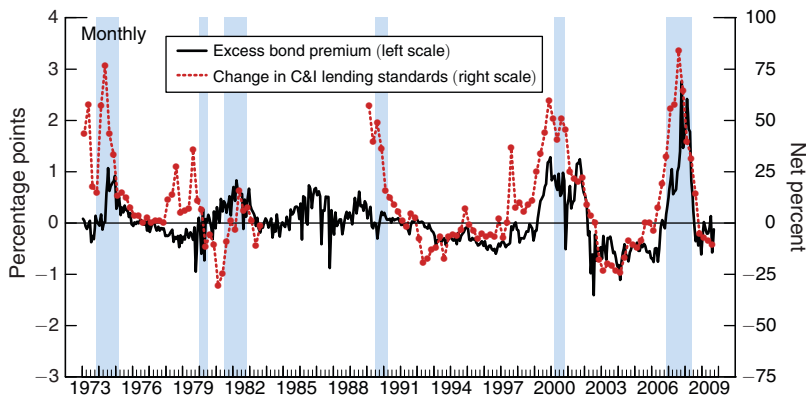
To the extent that financial disturbances alter risk perceptions in financial markets, changing risk attitudes of the marginal investors pricing corporate bonds may also influence a broader supply of credit. By and large, the corporate bond market is dominated by institutional investors such as large banks, insurance companies, and pension funds, intermediaries that possess specialized knowledge about the corporate bond market and in many cases are highly leveraged. These investors also face either explicit or implicit capital requirements, and as their financial capital becomes impaired, they act in a more risk-averse manner. This reduction in their effective risk-bearing capacity leads to an increase in the excess bond premium and a reduction in the supply of credit available to potential borrowers—both within the corporate cash market and through other sources of external finance—resulting in the type of asset market dynamics analyzed by He and Krishnamurthy (2010) and Adrian, Moench, and Shin (2010a, b).

Suggestive evidence supporting the link between the excess bond premium and risk attitudes and balance sheet conditions of financial intermediaries is provided in Figure 7. Panel A plots the excess bond premium against the diffusion index of the change in credit standards on commercial and industrial (C&I) loans at US commercial banks obtained from the Federal Reserve’s quarterly Senior Loan Officer Opinion Survey on Bank Lending Practices (SLOOS).¹³ The correlation between these two series—one obtained from a qualitative survey of commercial banks and the other obtained from market prices—is strikingly high, especially in the latter part of our sample.

Panel B highlights the link between the excess bond premium and the profitability of the US financial corporate sector as measured by its return on assets (ROA). Note that periods of elevated profitability are consistently associated with low levels of the excess bond premium, while the declines in ROA are mirrored by sharp increases in the excess bond premium. These comovements are consistent with the

¹³The SLOOS is usually conducted four times per year by the Federal Reserve Board, and up to 60 banks participate in each survey. Banks are asked to report whether they have changed their credit standards over the past three months on the major categories of loans to businesses and households. The series plotted is the net percentage of banks that reported tightening their credit standards on C&I loans to large and middle-market firms. Reported net percent equals the percent of banks that reported tightening their standards minus the percent that reported easing their standards. For the full text of the questions and more information on the survey, see <http://www.federalreserve.gov/boarddocs/SnLoanSurvey/>.

Panel A. Changes in bank lending standards and the excess bond premium



Panel B. Financial sector profitability and the excess bond premium

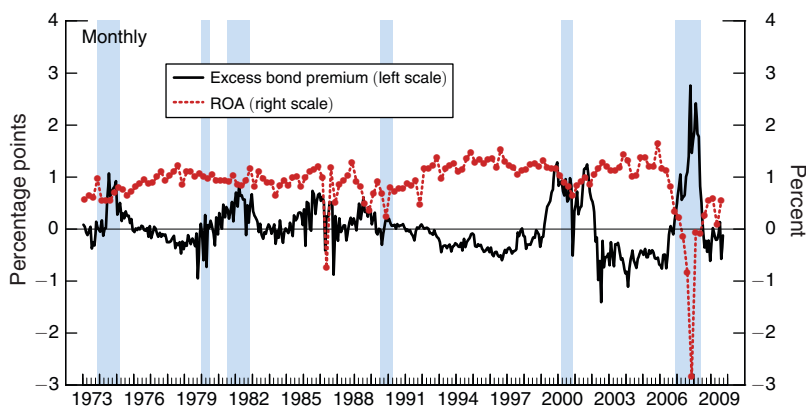


FIGURE 7. THE EXCESS BOND PREMIUM AND FINANCIAL MARKET CONDITIONS

Notes: Sample period: 1973:1–2010:9. The solid line in both panels depicts the estimated (option-adjusted) excess bond premium. The overlaid dots in panel A depict the net percent of SLOOS respondents that reported tightening their credit standards on C&I loans over the past three months. (There was no survey conducted during the 1984–89 period.) The overlaid dots in panel B depict the quarterly (annualized) ROA for the US financial corporate sector, calculated using Compustat data. The shaded vertical bars denote the NBER-dated recessions.

view that our proxy for the price of default risk responds to changes in the risk attitudes of financial intermediaries, at least as reflected in their willingness to make C&I loans and changes in the conditions of their balance sheets.

The 2007–2009 financial crisis offers a unique opportunity to explore this hypothesis further. Given that the origin of the crisis can undoubtedly be traced to the financial sector (e.g., Brunnermeier 2009 and Gorton 2009), we collected market-based data on the health of the financial sector, namely, the credit default swaps and equity valuations of primary dealers, major banks, and securities broker-dealers that trade in US Government securities with the Federal Reserve Bank of New York. By buying and selling an array of securities for a fee and holding an inventory of securities for resale, these highly leveraged financial intermediaries play a key role in most financial markets. As documented by Adrian and Shin (2010), broker-dealers differ from other

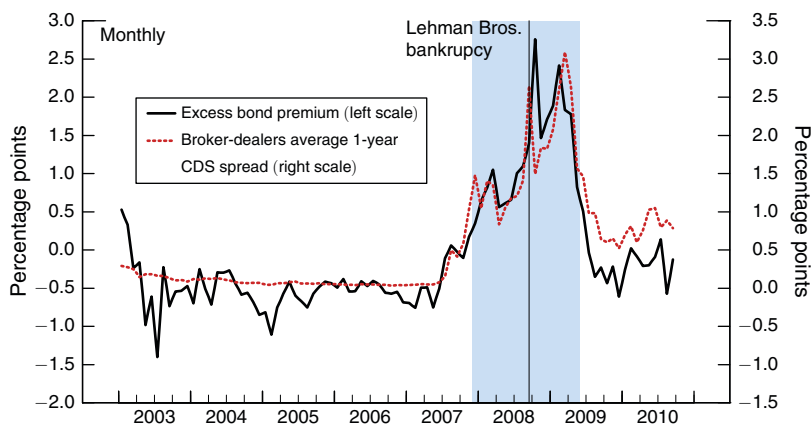


FIGURE 8. THE EXCESS BOND PREMIUM AND FINANCIAL INTERMEDIARY CDS SPREADS

Notes: Sample period: 2003:1–2010:9. The solid line depicts the estimated excess bond premium. The overlaid dotted line depicts the average one-year CDS spread of broker-dealers. The shaded vertical bar represents the 2007–09 NBER-dated recession.

types of institutional investors by their active procyclical management of leverage: expansions in broker-dealer assets are associated with increases in leverage as broker-dealers take advantage of greater balance sheet capacity; conversely, contractions in their asset holdings are associated with the deleveraging of their balance sheets.

The solid line in Figure 8 depicts the excess bond premium, while the overlaid dotted line represents the average one-year CDS spread for these institutions. The striking degree of comovement between the two series over the period shown again supports the interpretation that the excess bond premium fluctuates closely in response to movements in capital and balance sheet conditions of key financial intermediaries.¹⁴ Indeed, the collapse of Lehman Brothers on September 15, 2008—a watershed event in the recent crisis—provides a dramatic example of how disruptions in the effective risk-bearing capacity of the financial sector can influence the supply of credit.

To analyze more formally how shocks to the profitability of financial intermediaries affect our gauge of credit supply conditions, we consider a VAR, consisting of the option-implied volatility on the S&P 500 (VIX), the (value-weighted) excess market return, the (value-weighted) excess portfolio return of broker-dealers, the average one- and five-year broker-dealer CDS spreads, and the excess bond premium. By including both the one- and five-year CDS spreads, we allow such financial shocks to affect the market assessment of near- and longer-term default risk for these institutions. The VAR, using three lags of each endogenous variable, is estimated over the 2003:1–2010:9 period and also includes a dummy variable for September 2008.¹⁵

Within this multivariate framework, we trace out the impact of an orthogonalized shock to the excess return of broker-dealers, an innovation that, according to

¹⁴ Prior to 2003, only a small subset of broker-dealers had CDS contracts traded in the market.

¹⁵ Standard regression diagnostics revealed that this observation exerted an unduly large influence on the estimated coefficients.

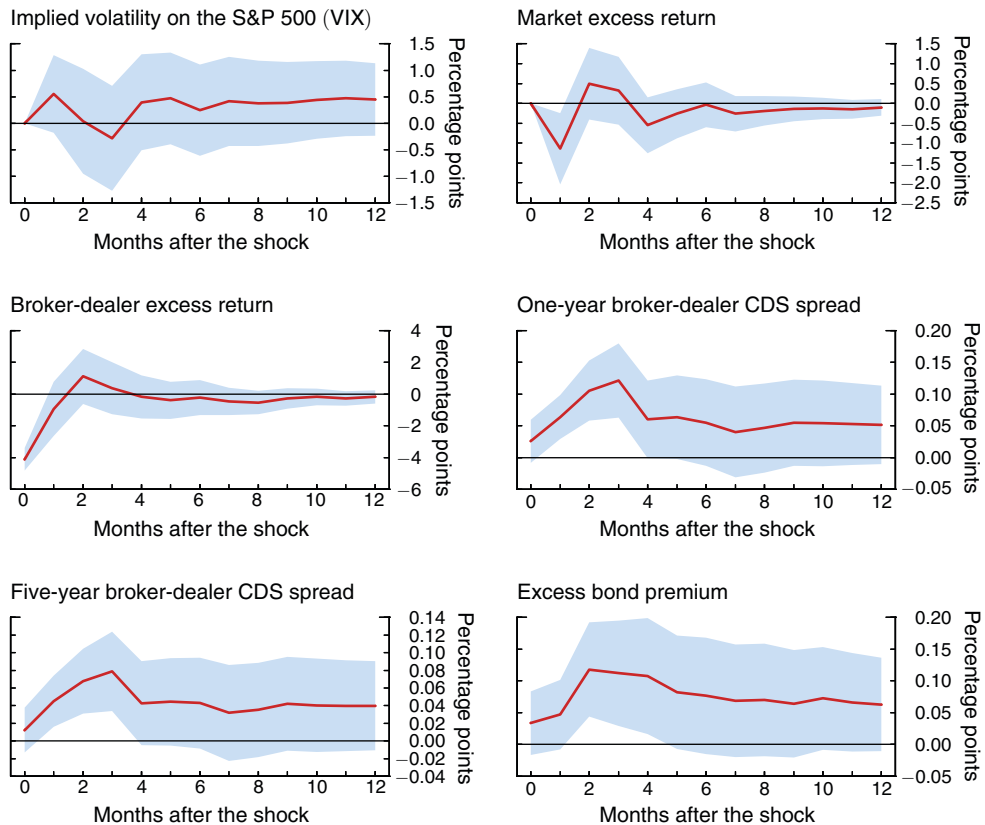


FIGURE 9. IMPLICATIONS OF A SHOCK TO THE PROFITABILITY OF FINANCIAL INTERMEDIARIES

Notes: The figure depicts the impulse responses to a one-standard-deviation orthogonalized shock to the average excess return of broker-dealers (see text for details). Shaded bands denote 95-percent confidence intervals based on 2,000 bootstrap replications.

our identification scheme, is uncorrelated with contemporaneous movements in the broad stock market and stock market volatility. The impulse responses shown in Figure 9 indicate that such an adverse shock to the profitability of these key financial intermediaries leads to an immediate rise in their near- and longer-term CDS spreads. Moreover, CDS spreads continue to widen for about three months after the initial impact, and they return only very gradually to their steady-state values. This persistent deterioration in investor assessment of the broker-dealers' creditworthiness is manifested by a sustained increase in the excess bond premium, the response of which is very close to that of the one-year CDS spread, likely the most accurate market-based indicator of near-term default risk in the financial sector.

Taken together, the evidence presented above is consistent with the view that systematic deviations in the pricing of corporate bonds relative to the expected default risk of the underlying issuer reflect shifts in the effective risk aversion of the financial sector. Increases in risk aversion lead to a decline in asset prices and a contraction in the supply of credit, both through the corporate bond market and the broader commercial banking sector, factors that contribute significantly to a resulting slow-down in economic activity.

V. Conclusion

This paper examined the role that credit spreads play in determining macroeconomic outcomes. We did so by constructing a new corporate bond credit spread index—the GZ credit spread—employing an extensive micro-level dataset of secondary market prices of outstanding senior unsecured bonds issued by a large panel of US nonfinancial corporations. Compared with the widely used default-risk indicators, the GZ credit spread was shown to be a robust predictor of future economic activity across a variety of economic indicators, short- and longer-term forecast horizons, and sample periods.

Using a flexible empirical framework, we then decomposed the GZ credit spread into two parts: a component reflecting the available firm-specific information on default risk and a residual component—the excess bond premium—that we argued likely represents variation in the pricing of default risk, rather than in the risk of default. According to our results, most of the predictive power of the GZ credit spread is accounted for by movements in the excess bond premium—indeed, over the 1985–2010 period, the excess bond premium accounts for all of the predictive content of the GZ credit spread for output growth.

Innovations to the excess bond premium that are orthogonal to the current state of the economy were shown to cause substantial and protracted contractions in economic activity, an appreciable disinflation, a decline in both short- and long-term risk-free rates, and a fall in the broad stock market. In turn, these shocks to the excess bond premium were linked to the deterioration in the profitability and creditworthiness of broker-dealers, marginal investors in the corporate debt market. All told, our findings are consistent with the notion that an increase in the excess bond premium reflects a reduction in the effective risk-bearing capacity of the financial sector and, as a result, a contraction in the supply of credit with significant adverse consequences for the macroeconomy.

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