Credit Spreads and Business Cycle Fluctuations

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April 9, 2011

Abstract

We re-examine the evidence on the relationship between credit spreads and economic activity, by constructing a credit spread index based on an extensive data set of prices of outstanding corporate bonds trading in the secondary market. Compared with the standard default-risk indicators, our credit spread index is a robust predictor of economic activity at both the short- and longer-term horizons. Using an empirical framework, we decompose the index into a predictable component that captures the available firm-specific information on expected defaults and a residual component—the excess bond premium—which we argue likely reflects variation in the price of default risk rather than variation in the risk of default. Our results indicate that the predictive content of credit spreads for economic activity is due primarily to movements in the excess bond premium. Innovations in the excess bond premium that are orthogonal to the current state of the economy are shown to lead to significant declines in economic activity and equity prices. We also show that a deterioration in the creditworthiness of broker-dealers—key financial intermediaries in the corporate cash market—causes an increase in the excess bond premium. These findings support the notion that a rise in the excess bond premium represents 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.

JEL Classification: E32, E44, G12
Keywords: corporate bond market, default-risk premium, financial intermediaries

We thank Viral Acharya, Eric Swanson, Min Wei, Jonathan Wright, and participants at the 2010 CEGE Conference on Financial Shocks and the Real Economy for helpful comments and suggestions. Robert Kurtzman and Michael Levere provided outstanding research assistance. All errors and omissions are our own responsibility alone. The views expressed in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or of anyone else associated with the Federal Reserve System.

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1 Introduction

Between the summer of 2007 and the spring of 2009, the U.S. economy was gripped by an acute liquidity and credit crunch, by all accounts, the most severe financial crisis since the Great Depression. At the height of the crisis in the autumn of 2008, the government, in an attempt to prevent the financial meltdown from engulfing the real economy, effectively assumed control of a number of systemically important financial institution; the Congress, faced with investors' rapidly deteriorating confidence in the financial sector, approved the plan to inject a massive amount of capital into the banking system; and the Federal Reserve dramatically expanded the number of emergency credit and liquidity facilities in an attempt to support the functioning of private debt markets.

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 financial theories that depart from the Modigliani and Miller [1958] paradigm of frictionless financial markets, 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 balance sheets 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 either case, credit spreads play a crucial role in the dynamic interaction of financial conditions with the real economy.

[1] Financial indicators considered in this vast literature include stock prices (Fama [1981] and Harvey [1989]); 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 et al. [2006]); spreads between rates on short-term commercial paper and rates on Treasury bills (Friedman and Kuttner [1992, 1998]; and Emery [1999]); and yield spreads on longer-term corporate debt (Gertler and Lown [1999]; Mody and Taylor [2004]; King et al. [2007]; Mueller [2007]; Gilchrist et al. [2009]; and Faust et al. [2011]).

In this paper, we re-examine the evidence on the relationship between corporate bond credit spreads and economic activity. To do so, we first construct a credit spread index—the "GZ credit spread"—that has considerable predictive power for economic activity over the 1973–2010 period. Our approach builds on the recent work of Gilchrist et al. [2009], in that we use prices of individual corporate bonds traded in the secondary market to construct this high-information content credit spread. According to our forecasting results, the predictive ability of the GZ credit spread for 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 resulting from a reduction in the expected present-value of corporate cash flows 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 default risk and credit spreads—the so-called excess bond premium.

The credit-spread 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 (e.g., Elton et al. [2001]). As shown by Collin-Dufresne et al. [2001], Houwelling et al. [2005], and Driessen [2005], the unexplained portion of the variation in credit spreads appears to reflect some combination of time-varying liquidity premium, to some extent the tax treatment of corporate bonds, and, most importantly for our purposes, a default-risk factor.

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 expected default risk of the issuer. These deviations, which most plausibly represent fluctuations in the price of default risk, have significant predictive ability for a wide range of economic activity indicators at both the near- and longer-term forecast horizons and over different sub-periods. This robust finding suggests that shifts in investor risk attitudes embedded in prices of corporate bonds are chiefly responsible for the considerable forecasting power of credit spreads for economic activity.

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, output, and appreciable disinflation. In addition, monetary policy is eased significantly in response to these adverse economic developments, and despite the decline in long-term Treasury yields, such shocks cause a sharp fall in the broad stock market.

To provide an interpretation for these "financial disruptions" in the context of the 2007–09 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 idiosyncratic shock to the equity valuations of these intermediaries leads to an immediate and persistent increase in their credit default swap (CDS) premiums, a response that is mirrored in 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. Consistent with the financial accelerator mechanisms emphasized by Kiyotaki and Moore [1997], Bernanke et al. [1999], and Hall [2010], this reduction in credit availability augurs a change in financial conditions with significant adverse consequences for macroeconomic outcomes.

2 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, the forward-looking nature of financial markets should cause the information about investors' expectations of future economic outcomes to become embedded in asset prices, though obtaining an accurate reading of this information can be complicated by the presence of time-varying risk premiums. As shown recently by Faust et al. [2011], credit spreads on long-term corporate bonds have been shown to be particularly useful for forecasting future economic activity, compared with a wide array of other asset market indicators

2.1 Data Sources and Methods

In this paper, we employ the "bottom-up" approach used by Gilchrist et al. [2009] to construct a credit spread index with a 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. Specifically, for a large sample of U.S. nonfinancial firms covered by the S&P's Compustat and the Center for Research in Security Prices (CRSP), month-end secondary market prices of their outstanding securities were obtained 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 senior unsecured issues with a fixed coupon schedule only.

The micro-level aspect of our data allows us to construct credit spreads that are not subject to the "duration mismatch" that plagues most commercially-available credit spread indexes. We do so by constructing for each individual corporate issue 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 $\{C(s) : s = 1, 2, \ldots, S\}$, 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 $\{C(s) : s = 1, 2, \ldots, S\}$ using continuously-compounded zero-coupon Treasury yields in period t, obtained from the U.S. Treasury yield curve estimated daily by Gürkaynak et al. [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.

[2] These two data sources include secondary market prices for a vast majority of dollar-denominated bonds publicly issued in the U.S. corporate cash market; see Gilchrist et al. [2009] for details.

Table 1: Summary Statistics of Corporate Bond Characteristics

Variable	Mean	SD	Min	P50	Max
No. of bonds per firm/month	2.91	3.64	1.00	2.00	74.0
Mkt. 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 (bps.)	204	281	5	118	3,499

Note: Sample period: 1973:M1–2010:M9; Obs. = 346,126; No. of bonds = 5,982; No. of firms = 1,112. Sample statistics are based on trimmed data (see text for details).
[a] Market value of the outstanding issue deflated by the CPI (1982–84 = 100).

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 with spreads 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 one 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 as many as 74 different senior unsecured bond issues trading in the market at a point in time.

The distribution of the real 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. However, because corporate bonds typically generate significant cash flows in the form of regular coupon payments, their duration is considerably shorter, with both the average and the median duration of a bit more than 6 years.

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 pre-specified conditions. Moreover, 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. However, over the past decade and a half, 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. Turning to returns, the (nominal) coupon rate on these bonds averaged 7.34 percent during our sample period, while the average nominal effective yield was 7.68 percent per annum. Reflecting the wide range of credit quality, the distribution of nominal yields is quite wide, with the minimum of 0.53 percent and the maximum of more than 44 percent. Relative to Treasuries, an average bond in our sample has an expected return of 204 basis points above the comparable risk-free rate, with the standard deviation of 281 basis points.

Using this micro-level data set, 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

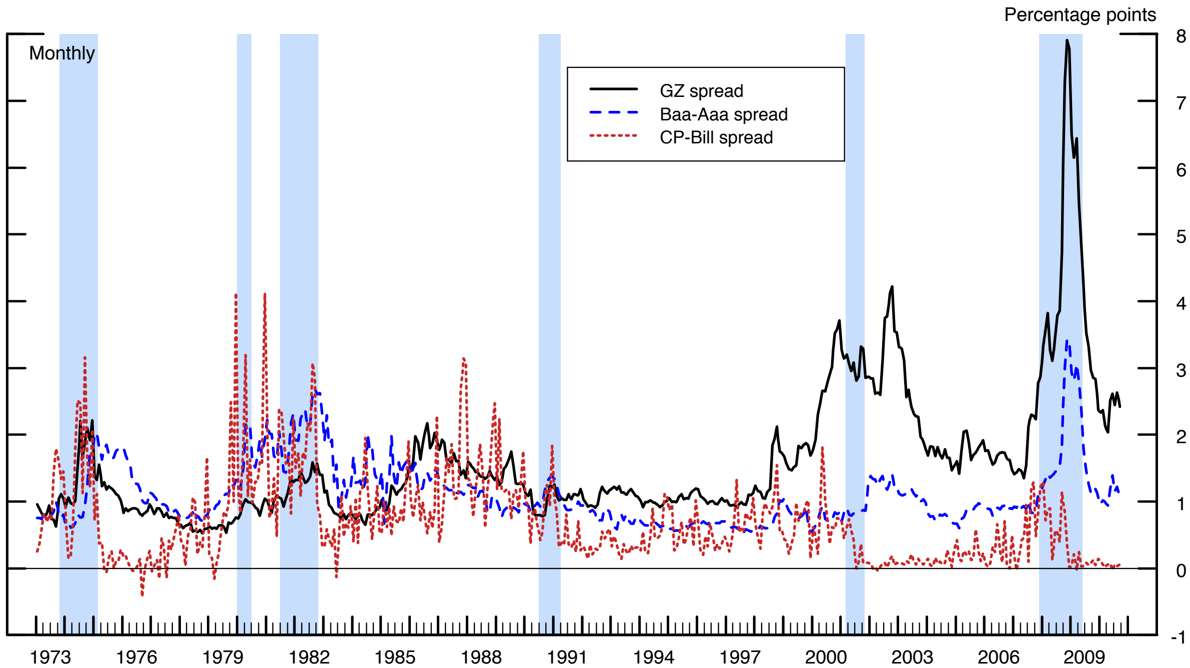
$$S_t^{GZ} = \frac{1}{N_t} \sum_i \sum_k S_{it}[k], \tag{1}$$

where N_t is the number of bond/firm observations in month t—that is, the GZ credit spread in month t is simply an arithmetic average of the credit spreads on outstanding bonds in that 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 yield spread between 1-month A1/P1-rated nonfinancial commercial paper and the 1-month Treasury yield (i.e., the paper-bill spread) and the spread between yields on indexes of Baa- and Aaa-rated seasoned industrial corporate bonds.[3]

All three credit spreads are clearly countercyclical, rising prior to and during economic

[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.

Figure 1: Selected Corporate Credit Spreads



NOTE: Sample period: 1973:M1–2010:M9. 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 1-month A1/P1 nonfinancial commercial paper and the 1-month Treasury yield (the dotted line). The shaded vertical bars represent the NBER-dated recessions.

downturns. Nonetheless, the pair-wise 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 spread is 0.21, whereas the paper-bill and the GZ spread are slightly negatively correlated, with the correlation coefficient of -0.17. Perhaps not too surprising, 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 the 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.

3 Credit Spreads and Economic Activity

This section examines the predictive power of the GZ credit spread for various measures of economic activity and compares its forecasting performance with that of several commonly-used financial indicators. Letting Y_t denote a measure of economic activity in period t, we define

$$\nabla^h Y_{t+h} \equiv \frac{c}{h+1} \ln\left(\frac{Y_{t+h}}{Y_{t-1}}\right); \quad (\nabla^0 = \nabla),$$

where $h \geq 0$ denotes 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).

To assess the predictive ability of credit spreads for future economic activity, we estimate the following univariate forecasting specification:

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

where 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 10-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 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]

Within this framework, we first analyze the information content of the three credit spreads

[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.

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.

3.1 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 and the real federal funds rate, along with p lags of ∇Y_{t-1}, as predictors. In line with the 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 signalling 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 the 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 point in the GZ credit spread in month t implies an almost 3.0 percentage points (annualized) drop

Table 2: Financial Indicators and Economic Activity

Forecast Horizon: 3 months

Financial Indicator	Payroll Employment				Unemployment Rate				Industrial Production			
Term spread	-0.096	-0.102	-0.110	-0.110	0.164	0.179	0.215	0.199	-0.182	-0.202	-0.239	-0.224
	[2.12]	[2.27]	[2.44]	[2.42]	[7.71]	[8.42]	[10.2]	[9.37]	[2.54]	[2.83]	[3.43]	[3.15]
Real FFR	-0.058	0.050	-0.038	-0.113	0.029	-0.152	-0.024	0.107	-0.035	0.183	0.016	-0.126
	[1.18]	[0.78]	[0.75]	[2.23]	[1.24]	[5.22]	[1.06]	[4.50]	[0.44]	[1.86]	[0.20]	[1.62]
CP–Bill spread	-	-0.165	-	-	-	0.268	-	-	-	-0.332	-	-
		[3.80]				[13.8]				[4.75]		
Baa–Aaa spread	-	-	-0.075	-	-	-	0.198	-	-	-	-0.211	-
			[2.05]				[10.4]				[3.08]	
GZ spread	-	-	-	-0.322	-	-	-	0.351	-	-	-	-0.386
				[8.50]				[19.5]				[5.28]
Adj. R^2	0.622	0.639	0.625	0.685	0.335	0.378	0.362	0.425	0.251	0.319	0.283	0.360

Forecast Horizon: 12 months

Financial Indicator	Payroll Employment				Unemployment Rate				Industrial Production			
Term spread	-0.252	-0.255	-0.241	-0.277	0.375	0.386	0.394	0.419	-0.358	-0.371	-0.357	-0.400
	[4.94]	[5.05]	[4.93]	[5.53]	[46.7]	[48.4]	[50.6]	[51.8]	[4.03]	[4.91]	[4.06]	[4.59]
Real FFR	-0.116	-0.064	-0.129	-0.204	0.037	-0.089	0.019	0.131	-0.094	0.052	-0.095	-0.175
	[2.10]	[0.96]	[2.45]	[3.84]	[4.60]	[9.72]	[2.45]	[16.4]	[0.98]	[0.48]	[1.02]	[1.90]
CP–Bill spread	-	-0.080	-	-	-	0.191	-	-	-	-0.226	-	-
		[2.29]				[36.0]				[3.67]		
Baa–Aaa spread	-	-	0.054	-	-	-	0.074	-	-	-	0.004	-
			[1.15]				[11.2]				[0.05]	
GZ spread	-	-	-	-0.497	-	-	-	0.453	-	-	-	-0.412
				[13.4]				[83.0]				[5.11]
Adj. R^2	0.422	0.424	0.422	0.579	0.270	0.292	0.273	0.417	0.227	0.258	0.225	0.346

NOTE: Sample period: 1973:M1–2010:M9. 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).

Table 3: Financial Indicators and Real GDP

Financial Indicator	Forecast Horizon: 1 quarter				Forecast Horizon: 4 quarters			
Term spread	-0.198	-0.217	-0.250	-0.247	-0.398	-0.406	-0.413	-0.460
	[1.77]	[1.92]	[2.07]	[2.26]	[2.79]	[2.81]	[2.70]	[3.22]
Real FFR	-0.016	0.175	0.020	-0.123	-0.036	0.042	-0.026	-0.131
	[0.12]	[1.12]	[0.15]	[0.95]	[0.24]	[0.22]	[0.17]	[0.87]
CP-bill spread	-	-0.254	-	-	-	-0.105	-	-
		[2.16]				[0.82]		
Baa–Aaa spread	-	-	-0.229	-	-	-	-0.066	-
			[1.95]				[0.52]	
GZ spread	-	-	-	-0.437	-	-	-	-0.482
				[4.96]				[5.74]
Adj. R^2	0.170	0.197	0.209	0.313	0.215	0.215	0.213	0.369

NOTE: Sample period: 1973:Q1–2010:Q3. 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).

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 two 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 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 signalling 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 four quarters.

4 The Excess Bond Premium

In this section, we exploit the micro-level aspect of our data to decompose the GZ credit spread into two components: a component that captures the systematic movements in default risk of individual firms and a residual component—the excess bond premium—which we argue most likely represents variation in the average price of bearing exposure to U.S. 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 $Z_{it}[k]$, according to

$$\ln S_{it}[k] = \beta DFT_{it} + \gamma' Z_{it}[k] + \epsilon_{it}[k]; \tag{3}$$

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 (cf. Cameron et al. [2011]).

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

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

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

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

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

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

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 $\widehat{S}_t^{GZ}$, versus movements in the excess bond premium EBP_t.

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

4.1 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 directly observed, 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 theoretically elegant and used widely by the financial industry, our choice of the Merton's framework is also motivated, in part, 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—can account quite well for the default-risk component of corporate bond prices. In particular, they show that such models generate *sensitivities* 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. Because in the contingent claims framework, any change in the value of debt is the result of a change in either the value of assets that collateralize the debt or in the riskless rate, their results imply that, to the extent that exposure to equity adequately reflects the underlying exposure to credit risk, this contingent claims approach is able to capture the default-related component of credit spreads.

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

$$dV = \mu_V V dt + \sigma_V V dW, \tag{4}$$

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:

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

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

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

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 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

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

Because 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

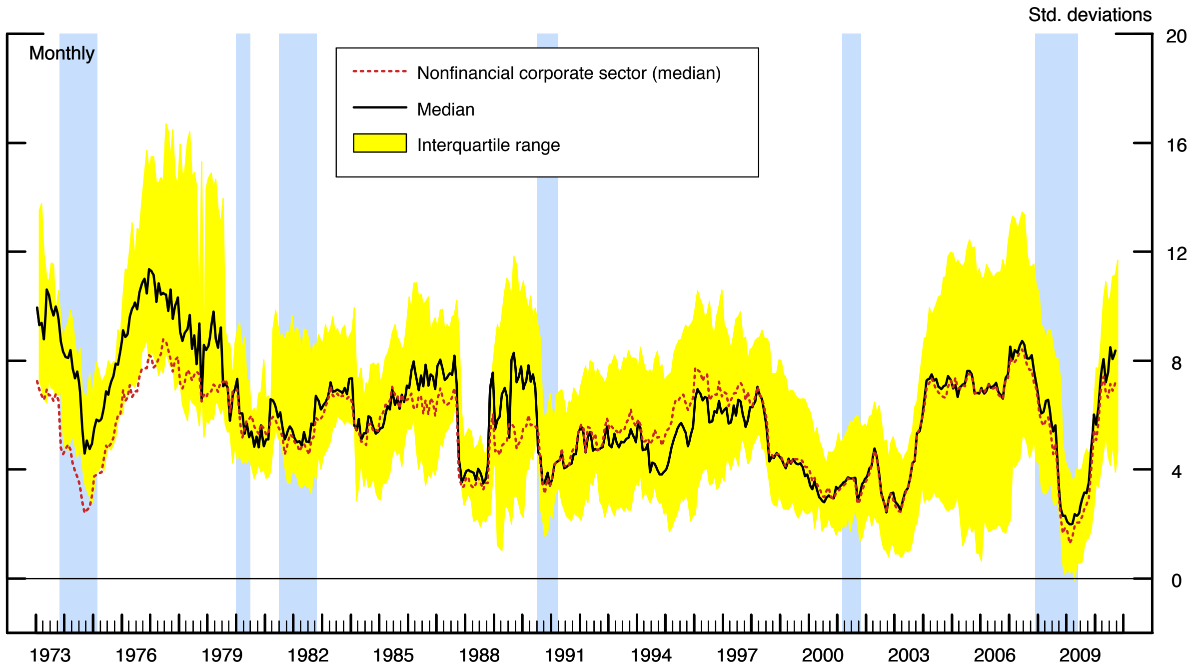
$$\sigma_E = \left[\frac{V}{E}\right] \Phi(\delta_1)\sigma_V. \tag{7}$$

From an operational standpoint, the most critical inputs to the Merton DD-model are clearly the market value of the equity E, the face value of the debt D, and the volatility of equity σ_E. Assuming a forecasting horizon of one year (i.e., $T = 1$), we implement the model in two 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 1-year constant-maturity Treasury yield), equations (5) and (7) can be solved for V and σ_V using standard numerical techniques.

As emphasized by Crosbie and Bohn [2003] and Vassalou and Xing [2004], 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

[7]This 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



NOTE: Sample period: 1973:M1–2010:M9. The figure depicts the distance-to-default (DD) calculated using the Merton [1974] model (see text for details). 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 U.S. nonfinancial corporate sector; all percentiles are weighted by the firm's outstanding liabilities. The shaded vertical bars represent the NBER-dated recessions.

of defaults and movements in financial asset prices. To resolve this problem, we implement an iterative procedure recently proposed by Bharath and Shumway [2008]. The procedure involves the following steps: First, we initialize the procedure by 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

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

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, distance-to-default measures by how many standard deviations must the log of this ratio 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 U.S. 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 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 pattern consistent with the jump in the GZ credit spread shown in Figure 1.

4.2 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 the bond-specific characteristics that could influence bond yields through either term or liquidity premiums. These (pre-determined) characteristics—denoted by the vector $Z_{it}[k]$—include the bond's duration ($DUR_{it}[k]$), the amount outstanding ($PAR_{it}[k]$), the (fixed) coupon rate ($CPN_i[k]$), the age of the issue ($AGE_{it}[k]$), and an indicator variable that equals one if the bond is callable and zero otherwise ($CALL_i[k]$).

The regression also includes industry (3-digit NAICS) fixed effects to control for any systematic (time-invariant) differences in expected recovery rates across industries. Lastly, the specification includes credit rating (S&P) 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 (e.g., Löffler [2004, 2007]).[9]

As shown in Table 4, our market-based measure of default risk is statistically a highly

[8] We eliminated from our sample all observations with the DD of more than 20 or less than -2, cutoffs corresponding roughly to the 99th and 1st percentiles of the DD distribution, respectively.

[9] We 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 (e.g., Levin et al. [2004]). The inclusion of these terms, however, had virtually no effect on any of our results.

Table 4: Credit Spreads and the Distance-to-Default

Explanatory Variable	*Est.*	*S.E.*	*Est.*	*S.E.*
$-DD_{it}$	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
$\ln(CPN_i[k])$	0.439	0.074	0.031	0.062
$\ln(AGE_{it}[k])$	0.047	0.008	0.135	0.010
$CALL_i[k]$	0.262	0.029	-0.427	0.210
$-DD_{it} \times CALL_i[k]$	-	-	-0.030	0.004
$\ln(DUR_{it}[k]) \times CALL_i[k]$	-	-	-0.120	0.023
$\ln(PAR_{it}[k]) \times CALL_i[k]$	-	-	-0.122	0.024
$\ln(CPN_i[k]) \times CALL_i[k]$	-	-	0.915	0.078
$\ln(AGE_{it}[k]) \times CALL_i[k]$	-	-	-0.132	0.013
$LEV_t \times CALL_i[k]$	-	-	-0.385	0.027
$SLP_t \times CALL_i[k]$	-	-	-0.088	0.017
$CRV_t \times CALL_i[k]$	-	-	-0.041	0.019
$VOL_t \times CALL_i[k]$	-	-	0.134	0.021
Adj. R^2	0.649		0.700	
Industry Effects[a]	0.000		0.000	
Credit Rating Effects[b]	0.000		0.000	

NOTE: Sample period: 1973:M1–2010:M9. Obs. = 346,126; No. of bonds/firms = 5,982/1,112. Dependent variable is $\ln(S_{it}[k])$, 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 10-year Treasury yield. Asymptotic standard errors are clustered in both the firm (i) and time (t) dimensions, according to Cameron et al. [2011].
[a] p-value of the exclusion test of industry fixed effects.
[b] p-value of the exclusion test of credit rating fixed effects.

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.

According to the Merton model, the distance-to-default should summarize all available information regarding the risk of default. Consequently, movements in the risk-free interest rates should affect credit spreads only insofar that they change the 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 the 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 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.

One possible way to deal with this issue would be to confine the analysis to a sub-sample of noncallable bonds. However, as reported in Table 1, callable bonds account, on average, for two-third 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 the 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 $Z_{it}[k]$ in equation (3) with the $CALL_i[k]$ 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 be also affected by the realized monthly volatility of the daily 10-year Treasury yield, a proxy for interest rate uncertainty.[10]

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 the 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 the 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

[10] The level, slope, and curvature factors correspond, respectively, to the first three principal components of nominal Treasury yields at 3-month, 6-month, 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 3- and 6-month bill rates are derived from the smoothed Treasury yield curve estimated by Gürkaynak et al. [2007].

Table 5: Selected Marginal Effects by Type of Bond

	Noncallable		Callable			
Explanatory Variable	*Est.*	*S.E.*	*Est.*	*S.E.*	*Mean*[a]	*STD*[b]
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

Note: 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 the mean equal to zero and the 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.

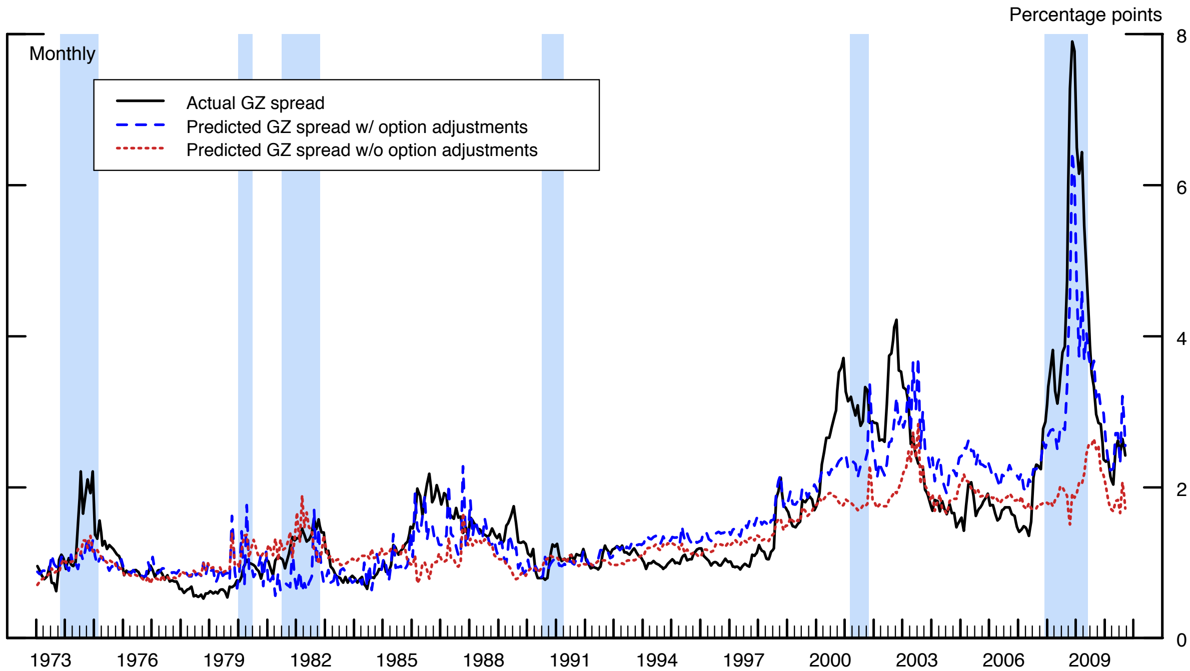
a 13 basis points rise in the spreads of their callable counterparts.

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 the 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 and one that does not. Over most of our sample period, the option adjustment 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 a more volatile fitted values.

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 a historically low level. 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

Figure 3: Actual and Predicted Credit Spreads



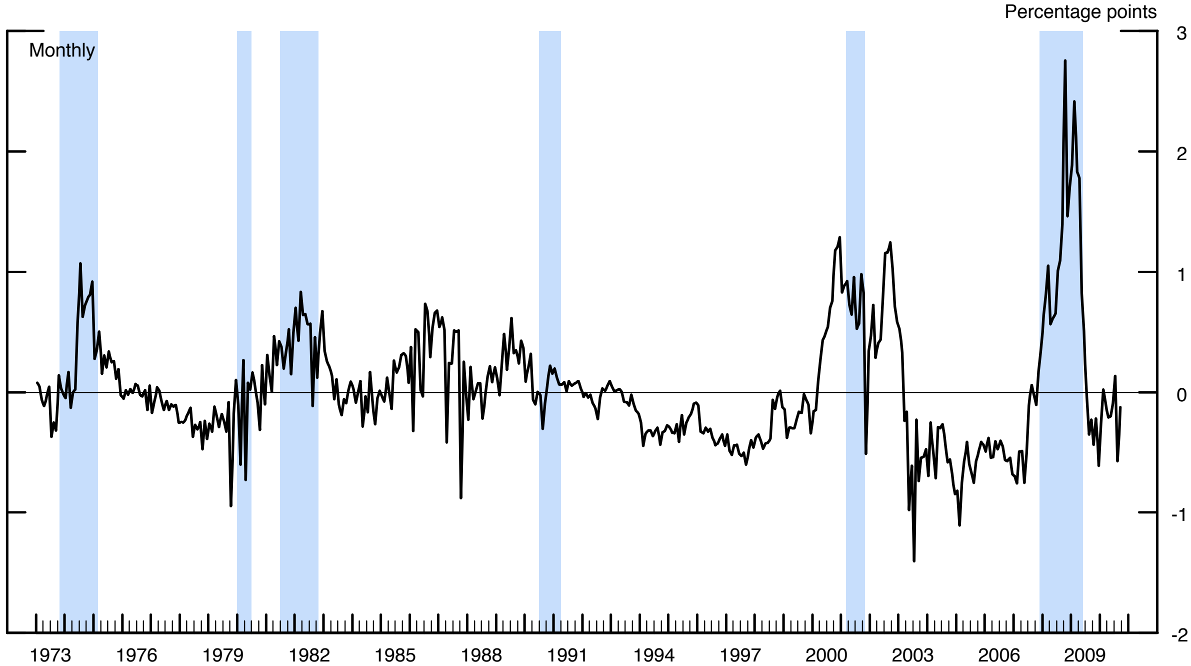
NOTE: Sample period: 1973:M1–2010:M9. The solid line depicts the actual GZ credit spread. The dashed line depicts the predicted GZ credit spread based on the specification that includes the term structure option-adjustment terms; the dotted line depicts the predicted GZ credit spread based on the specification that excludes the term structure option-adjustment terms (see text for details). The shaded vertical bars represent the NBER-dated recessions.

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 1990–91 recession, the premium increased significantly 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 U.S. 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

Figure 4: The Excess Bond Premium



NOTE: Sample period: 1973:M1–2010:M9. The figure depicts the estimated (option-adjusted) excess bond premium. The shaded vertical bars represent the NBER-dated recessions.

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.

5 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, movements that arguably reflect variation in the pricing of default risk rather than variation in the risk of default. We now examine whether movements in the excess bond premium provides independent information about future economic activity. First, we analyze the extent to which the forecasting power of the GZ credit spread documented in Section 3 is attributable to the predicted component ($\widehat{S}_t^{GZ}$), versus the excess bond premium (EBP_t).

Table 6: The Excess Bond Premium and Economic Activity

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

NOTE: Sample period: 1973:M1–2010:M9. 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).
[a] Excluding the estimated call-option adjustment effect.

We then add the excess bond premium to an otherwise standard macroeconomic VAR and examine the implications of the orthogonalized shocks to the excess bond premium for the real economy and asset prices more generally.

5.1 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—$\widehat{S}_t^{GZ}$ and EBP_t—are allowed to enter the forecasting regression (2) separately. According to our estimates, both the excess bond premium and the predicted GZ credit spread contains significant independent explanatory power for all three economic indicators, at both the 3- and 12-month forecast horizons. However, the (absolute) magnitude of the estimated coefficients on the excess bond premium tends to be significantly larger than that of the coefficients associated with predicted GZ spread, a finding indicating that the information content of credit spreads for economic activity largely reflects fluctuations in the non-default 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. For brevity, we report the results for the year-ahead forecast horizon only. We do, however, perform an important robustness check by performing the analysis for the 1985–2010 subsample, a period marked by significant deregulation of financial markets and changes in the conduct of monetary policy. In combination with the rapid advances in information technology, these changes in financial landscape have lowered the information and monitoring costs of investments in public securities and have increased the tendency of corporate borrowing to take the form of negotiable securities issued directly in capital markets, thereby improving liquidity in both the primary and secondary corporate debt markets.

As indicated in the first column of the top panel 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 four 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 points increase leads to 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 investment and spending on 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 the bottom panel, 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

Table 7: The Excess Bond Premium, Real GDP, and its Main Components

Sample Period: 1973:Q1–2010:Q3 (Forecast Horizon: 4 quarters)								
Financial Indicator	GDP	C-NDS	C-D	I-RES	I-ES	I-HT	I-NRS	INV
Term spread	-0.478	-0.452	-0.551	-0.564	-0.398	-0.098	0.317	-0.123
	[3.33]	[3.89]	[2.55]	[5.23]	[3.16]	[0.83]	[2.73]	[1.43]
Real FFR	-0.036	0.106	0.106	-0.003	-0.086	-0.092	-0.111	0.014
	[0.24]	[0.99]	[0.58]	[0.03]	[0.82]	[0.67]	[0.87]	[0.15]
Predicted GZ spread[a]	-0.258	-0.209	0.014	-0.159	-0.221	-0.426	-0.186	-0.287
	[2.56]	[2.39]	[0.11]	[2.10]	[2.48]	[4.43]	[2.07]	[4.11]
Excess bond premium	-0.364	-0.260	-0.127	-0.018	-0.558	-0.374	-0.587	-0.656
	[5.36]	[4.36]	[1.00]	[0.29]	[5.87]	[4.42]	[5.77]	[9.39]
Adj. R^2	0.365	0.349	0.224	0.419	0.481	0.432	0.557	0.580
Sample Period: 1985:Q1–2010:Q3 (Forecast Horizon: 4 quarters)								
Financial Indicator	GDP	C-NDS	C-D	I-RES	I-ES	I-HT	I-NRS	INV
Term spread	-0.509	-0.362	-0.456	-0.596	-0.340	-0.071	0.392	-0.321
	[4.09]	[4.07]	[2.00]	[6.52]	[2.75]	[0.60]	[2.46]	[4.08]
Real FFR	0.424	0.181	0.395	0.331	0.032	-0.130	-0.000	0.301
	[2.67]	[1.53]	[1.49]	[2.96]	[0.18]	[0.90]	[0.00]	[2.47]
Predicted GZ spread[a]	-0.023	-0.093	0.194	0.045	-0.088	-0.294	-0.061	-0.047
	[0.20]	[0.82]	[0.99]	[0.47]	[0.63]	[2.11]	[0.49]	[0.46]
Excess bond premium	-0.501	-0.362	-0.260	-0.035	-0.650	-0.382	-0.613	-0.701
	[6.80]	[5.02]	[1.51]	[0.50]	[5.41]	[3.39]	[4.76]	[8.20]
Adj. R^2	0.357	0.508	0.101	0.484	0.448	0.405	0.624	0.635

NOTE: Dependent variable is $\nabla^4 Y_{t+4}$, where Y_t denotes real GDP or one of its components in quarter t: C-D = PCE on durable goods; C-NDS = PCE on nondurable goods & services; I-RES = residential investment; I-ES = business fixed investment in E&S (excl. 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).

[a] Excluding the estimated call-option adjustment effect.

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 points increase in the excess bond premium in quarter t lowers output about 2.0 percentage points over the next four quarters.

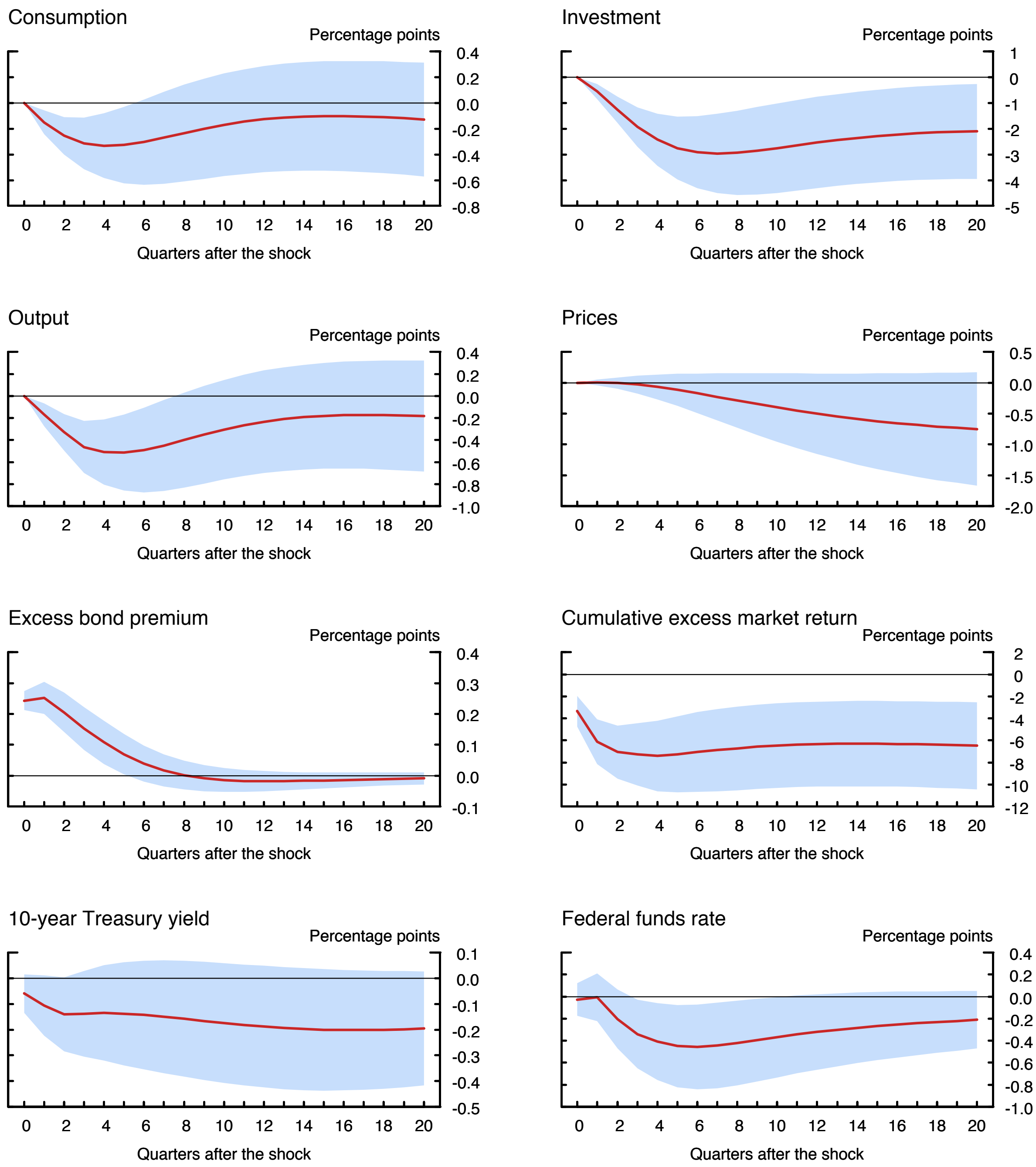
In summary, the above analysis indicates that the excess bond premium is a robust predictor of economic activity. This finding holds true across a variety of economic indicators, short- and longer-term forecast horizons, and sample periods. Furthermore, our forecasting results imply that since the mid-1980s,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

5.2 Macroeconomic Implications

In this section, we examine macroeconomic consequences of shocks to the excess bond premium. We do so by adding the premium to a standard VAR that includes the following endogenous variables: (1) log-difference of real personal consumption expenditures (PCE); (2) log-difference of real business fixed investment (BFI); (3) log-difference of real GDP; (4) inflation as measured by the log-difference of the GDP price deflator; (5) the quarterly average of the excess bond premium; (6) the quarterly value-weighted excess stock market return from CRSP; (7) the 10-year (nominal) Treasury yield; and (8) 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 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.

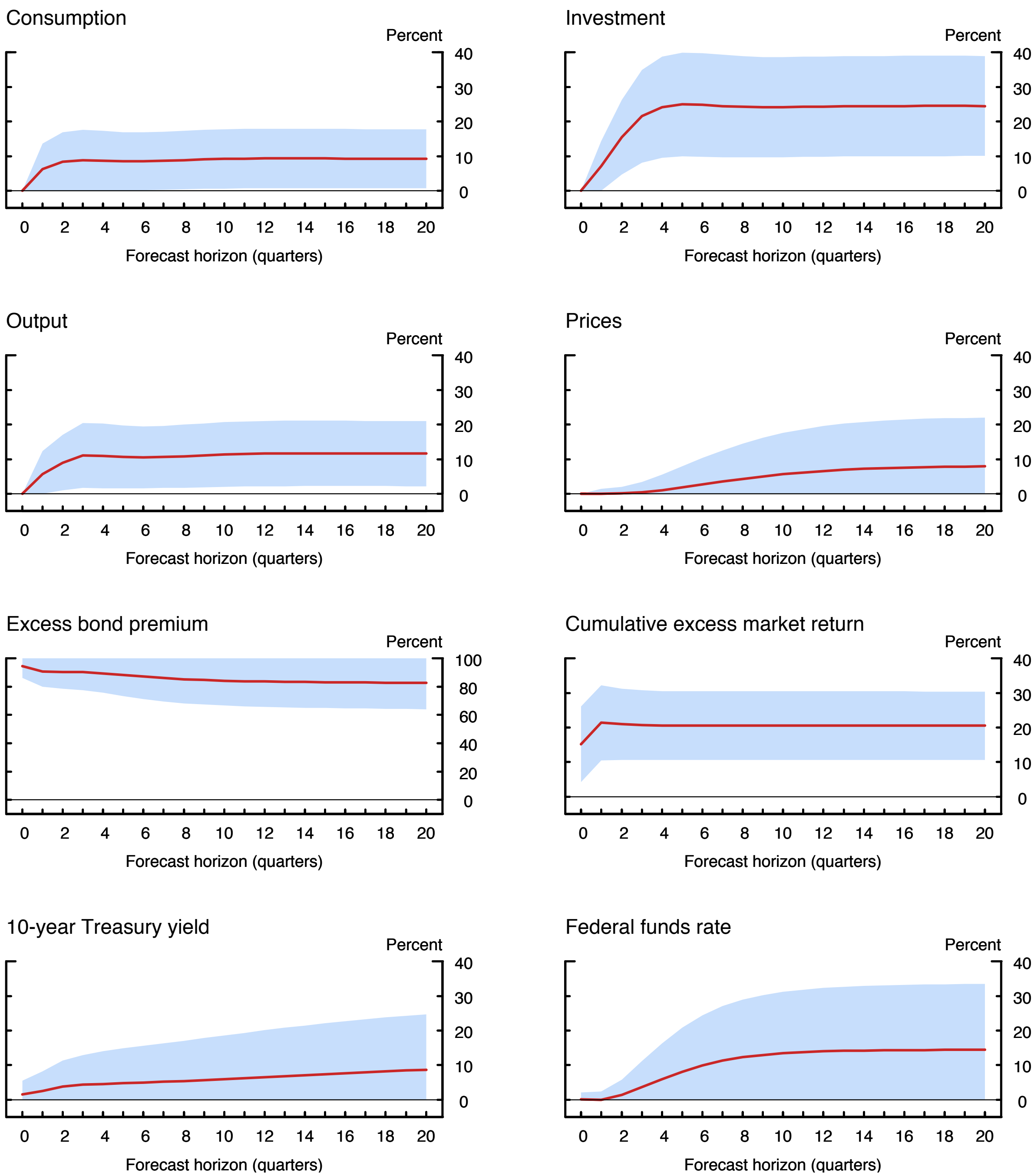
Figure 5 depicts the impulse response functions of the endogenous variables to an orthogonalized shock to the excess bond premium. An unanticipated increase of one standard deviation in the excess bond premium—about 20 basis points—causes a significant reduction in real economic activity, with consumption, 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.

Figure 5: Macroeconomic Implications of a Shock to the Excess Bond Premium



Note: The figure depicts the impulse responses to a negative 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: Forecast Error Variance Decomposition



NOTE: 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.

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 the cumulative decline of about 7.0 percentage points relative to trend growth.

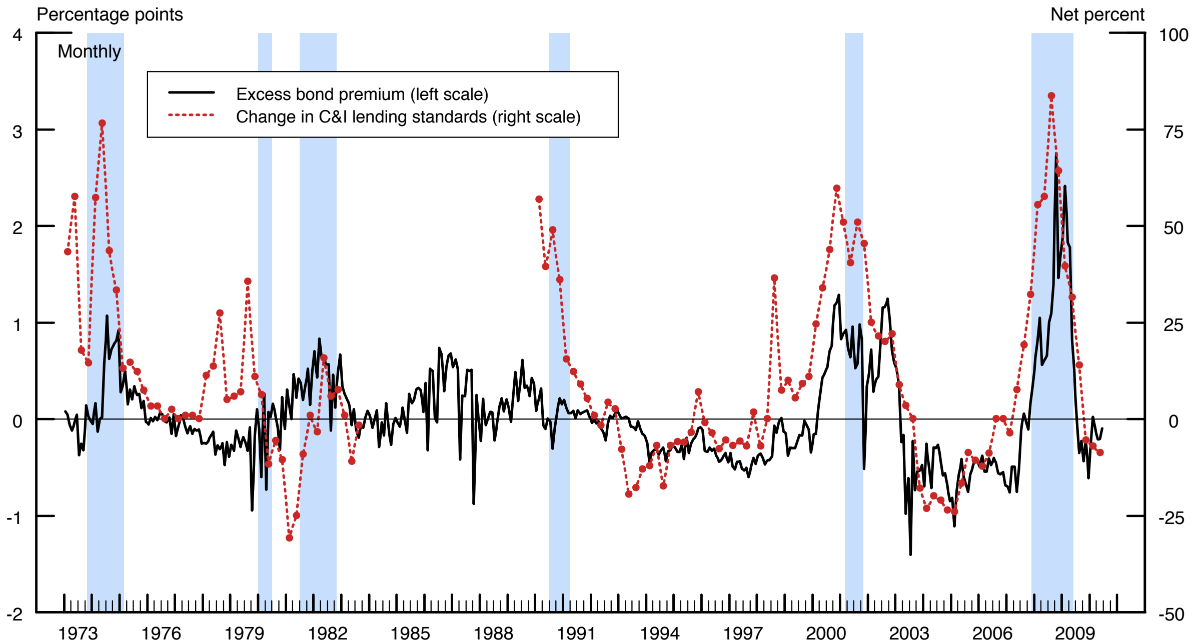
Figure 6 shows the amount of variation in the endogenous variables explained by the orthogonalized shocks to the excess bond premium. These financial disruptions account for more than 10 percent of the variation in output and 25 percent of the variation in investment at business cycle frequencies, proportions 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 equity prices.

These results 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 Kiyotaki and Moore [1997], Bernanke et al. [1999], and Hall [2010]. Our findings also provide empirical support for the recent work of Gertler and Kiyotaki [2009] and Gertler and Karadi [2010], 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.

To the extent that financial shocks cause variation in the risk attitudes of the marginal investors pricing corporate bonds, changing risk attitudes of these intermediaries may also influence the supply of credit available through the corporate bond market. By and large, the corporate bond market is dominated by institutional investors such as large banks, insurance companies, and pensions funds, intermediaries that possess specialized knowledge about the corporate bond market and are often 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 banking system and to those who rely on the corporate cash market—resulting in the type of asset market dynamics analyzed by He and Krishnamurthy [2010] and Adrian et al. [2010a,b].

A piece of evidence in favor of such credit-supply mechanisms is provided in Figure 7,

Figure 7: The Excess Bond Premium and Changes in Banks' Credit Standards

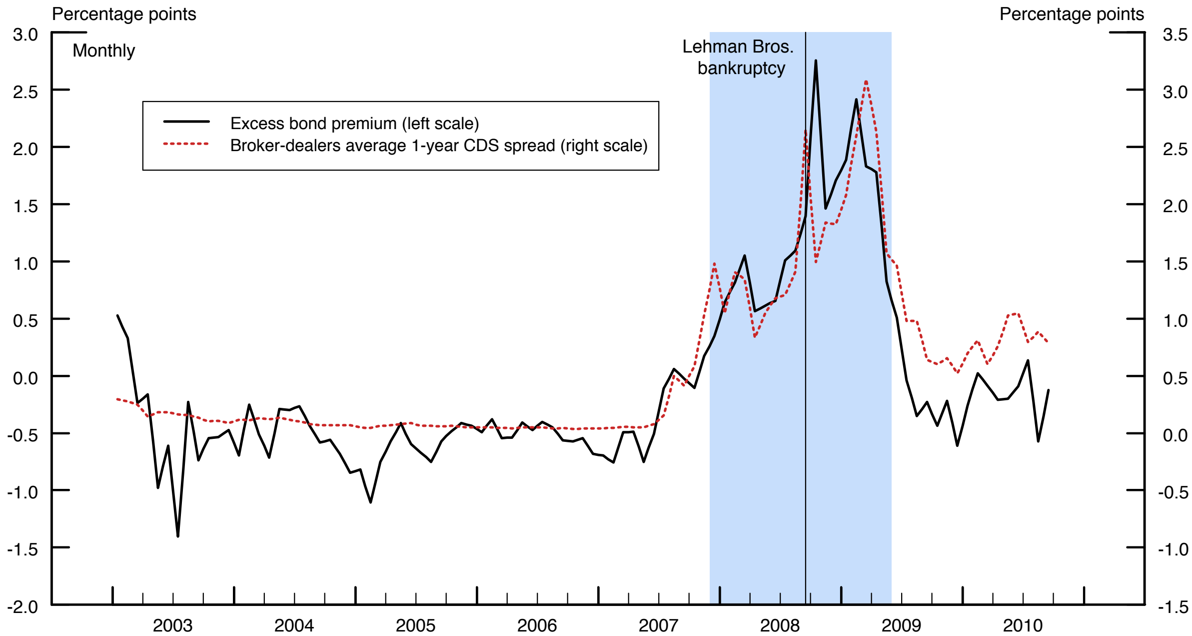


NOTE: Sample period: 1973:M1–2010:M9. The solid line depicts the estimated excess bond premium. The overlayed dots depict the net percent of respondents to the Federal Reserve's quarterly Senior Loan Officer Opinion Survey on Bank Lending Practices that reported tightening their credit standards on C&I loans to large and middle-market firms over the past three months. Reported net percent equals the percent of banks that reported tightening their standards minus the percent that reported easing their standards. (There was no survey conducted during the 1984-89 period.) The shaded vertical bars denote NBER-dated recessions.

which plots the excess bond premium against the diffusion index of the change in credit standards on commercial and industrial (C&I) loans at U.S. commercial banks obtained from the Federal Reserve's quarterly Senior Loan Officer Opinion Survey on Bank Lending Practices. 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. In effect, the willingness of banks to make business loans comoves strongly with the supply conditions in corporate cash market as measured by the excess bond premium.

To explore this hypothesis further and to shed some light on the 2007–09 financial crisis, we collected data on credit default swaps and equity valuations of primary dealers—that is, major banks and securities broker-dealers that trade in U.S. 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 intermedi-

Figure 8: The Excess Bond Premium and Financial Intermediary CDS Spreads



NOTE: Sample period: 2003:M1–2010:M9. The solid line depicts the estimated excess bond premium. The overlayed dotted line depicts the average 1-year CDS spread of broker-dealers. The shaded vertical bar represents the 2007–09 NBER-dated recession.

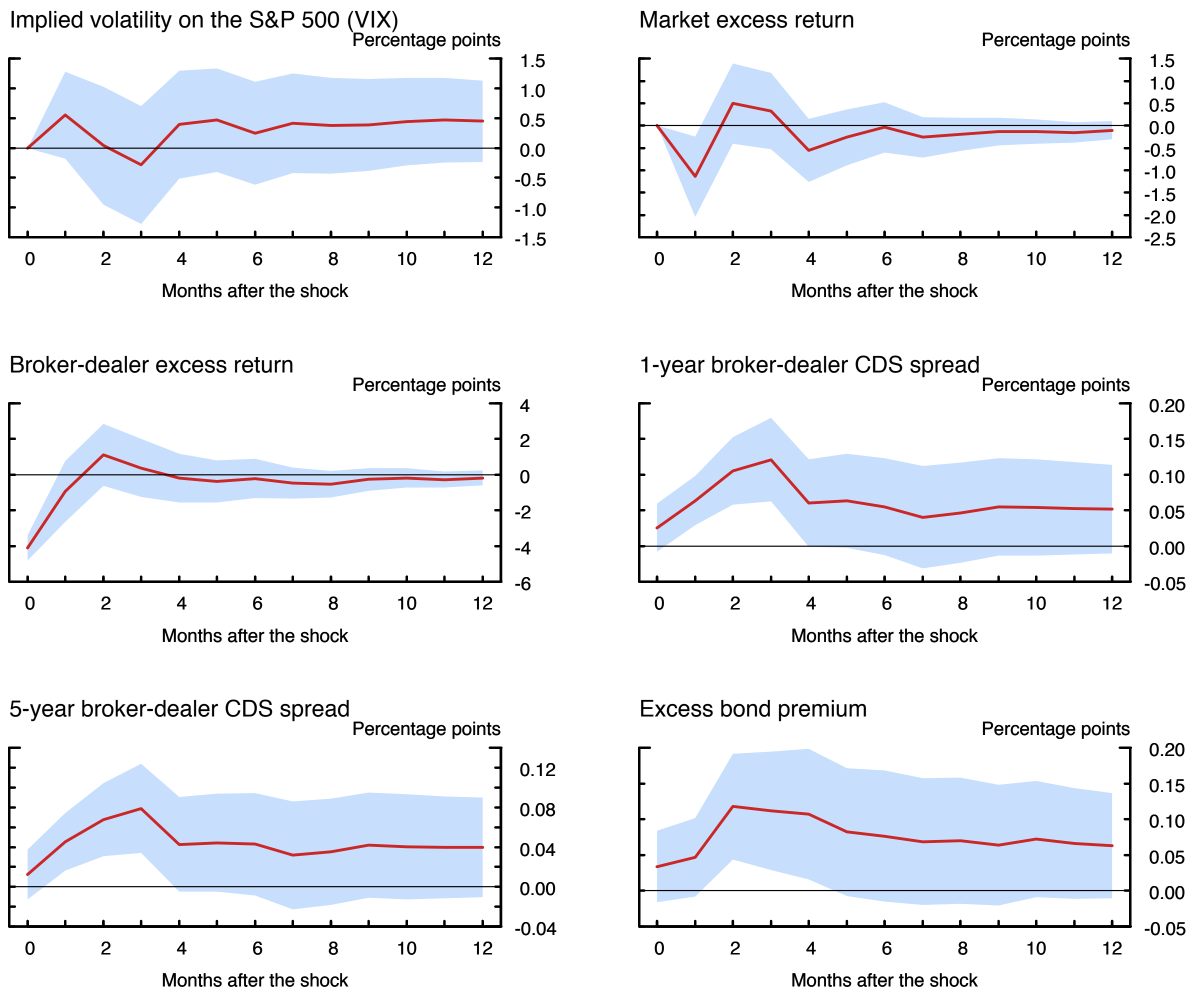
aries play a key role in most financial markets. As documented by Adrian and Shin [2010], broker-dealers differ from other types of institutional investors by their active pro-cyclical management of leverage: Expansions in broker-dealer assets are associated with increases in leverage as broker-dealers take advantage of greater balance sheet capacity; conversely, contractions in their assets are associated with the de-leveraging of their balance sheets.

The solid line in Figure 8 depicts the excess bond premium, while the overlayed solid line represents the average 1-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—our proxy for the price of default risk—fluctuates closely in response to movements in capital and balance sheet conditions of financial intermediaries.[11] Indeed, the collapse of Lehman Brothers on September 15, 2008—a watershed event in the history of 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 how shocks to the profitability of financial intermediaries affect our gauge of

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

Figure 9: Implications of a Shock to the Profitability of Financial Intermediaries



Note: The figure depicts the impulse responses to a negative 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.

credit-supply conditions, we estimate 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 1- and 5-year broker-dealer CDS spreads, and the excess bond premium. The VAR, using two lags of each endogenous variable, is estimated over the 2003:M1–2010:M9 period and also includes a dummy variable for September 2008.[12]

[12]Standard regression diagnostics revealed that this observation exerted an unduly large influence on the estimated coefficients. By dummying out the bankruptcy of Lehman Brothers, we are ensuring that our results are not driven by a single extreme observations, an important consideration given the relatively short

Within this framework, we trace out the impact of an orthogonalized shock to the excess return of broker-dealers, a shock that, according to our identification scheme, is uncorrelated with the contemporaneous movements in the broad stock market and stock market volatility. The impulse responses shown in Figure 9 indicate that such an adverse idiosyncratic 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 the investor assessment of the broke-dealers' creditworthiness is manifested by a sustained increase in the excess bond premium, the response of which is very close to that of the 1-year CDS spread, likely the most accurate indicator of very near-term default risk in the financial sector.

6 Conclusion

This paper re-examined the role that corporate bond 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 data set of secondary market prices of outstanding senior unsecured bonds for a large panel of U.S. nonfinancial corporation. 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 indicator and forecast horizons.

Using an 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 plausibly captured variation in the pricing of default risk. 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.

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, the marginal

sample. Nonetheless, our results are robust to the exclusion of the September 2008 dummy from the VAR.

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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