11TH JACQUES POLAK ANNUAL RESEARCH CONFERENCE
NOVEMBER 4–5, 2010

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Paper presented at the 11th Jacques Polak Annual Research Conference
Hosted by the International Monetary Fund
Washington, DC─November 4–5, 2010

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October 27, 2010

Abstract

The depth and duration of the 2007–09 economic downturn serves as a powerful reminder of the real consequences of financial shocks. Although channels through which disruptions in financial markets can affect economic activity are relatively well understood from a theoretical perspective, assessing their quantitative implications for the real economy remains a considerable challenge. This paper examines the extent to which the workhorse New Keynesian model—augmented with the standard financial accelerator mechanism—is capable of producing the dynamics of the U.S. economy during the recent financial crisis. To do so, we utilize secondary market prices of outstanding bonds of U.S. financial institutions to construct a measure of financial shocks, which are then used to simulate the model over the crisis period. Our results indicate that a reasonably calibrated version of the model can closely match the observed declines in consumption, investment, output, and hours worked; in addition, the model can account for the sharp widening of nonfinancial credit spreads, a decline in nominal short-term interest rates and for the persistent disinflation experienced in the wake of financial disruptions. Given its empirical relevance, we then use this framework to analyze the potential benefits of a monetary policy rule that allows the short-term nominal rate to respond to changes in financial conditions as measured by the movements in credit spreads. Our results indicate that such a spread-augmented policy rule can effectively dampen the negative consequences of financial disruptions on real economic activity, while engendering only a modest increase in inflation.

JEL Classification: E30, E44, E52

Keywords: financial disruptions, economic activity, spread-augmented monetary policy rule

This paper was prepared for the Eleventh Jacques Polak Annual Research Conference on "Macroeconomic and Financial Policies after the Crisis." We thank Ed Nelson for helpful comments. Michael Levere provided outstanding research assistance. 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

The complexity and sophistication of today's financial instruments and institutions—in a global economy with a high degree of financial integration—were undoubtedly the major factors behind the extraordinarily rapid transmission of financial shocks during the recent crisis. When rising delinquencies on subprime mortgages in the first half of 2007, triggered by the end of the housing boom in the United States, started to lead to large losses on related structured credit products, investors became greatly concerned about structures of securitized financial products more generally and began to pull back from risk-taking. In the late summer, with investors' risk appetite diminished substantially, the short-term funding markets in the United States and abroad became severely disrupted, and liquidity in private credit markets dropped sharply.

The initial financial turmoil did not appear to leave much of an imprint on real economic activity. However, the persistent and escalating pressures on bank balance sheets caused a pronounced tightening of aggregate credit conditions, a drop in asset values, and a slump in business and consumer confidence. Indeed, on December 1, 2008, the NBER's Business Cycle Dating Committee determined that a peak in U.S. economic activity occurred sometime in December 2007. And in spite of a number of unprecedented policy actions by the Federal Reserve and other U.S. government entities to arrest and mitigate the ensuing contraction in economic activity, the 2007–09 downturn has entered the record as the most severe—in terms of both its depth and duration—recession of the postwar period.

The destructive power of this "adverse feedback loop" between financial conditions and the real economy has led to much soulsearching among policymakers and economists. The debate among the former, in particular, has focused on whether central banks should respond only to inflation in the price of goods and economic slack, or should they also respond to movements in asset prices. The latter group, in contrast, has responded by developing a slew of new dynamic general equilibrium models, in which the deterioration in the equity capital position, or net worth, of financial intermediaries—by reducing the supply of credit—leads to and amplifies the ensuing economic downturn.[1]

Although channels through which disruptions in financial markets can influence eco-

[1] Empirical studies documenting the real-side effects of adverse credit supply shocks include Peek and Rosengren [1997, 2000], Calomiris and Mason [2003], and Ashcraft [2005]. From a theoretical perspective, Goodfriend and McCallum [2007] investigate the role of banks that produce loans and deposits using a production function that requires, as inputs, both the monitoring effort and collateral; Van den Heuvel [2008] analyzes the welfare costs of regulatory capital requirements, which reduce the ability of banks to create liquidity; Dib [2009] and Gerali et al. [2010] formulate DSGE models with monopolistically competitive banks in deposits and loan markets; and Gertler and Karadi [2009] construct a model in which an agency problem between depositors and financial intermediaries ties the availability of credit to intermediaries' capital position. An alternative approach, followed by Chen [2001], Meh and Moran [2004], and Hirakata et al. [2009], has been to incorporate the Holmstrom and Tirole [1997] framework into quantitative general equilibrium models.

nomic activity are relatively well understood from a theoretical perspective, assessing their quantitative implications for the real economy remains a considerable challenge. In this paper, we examine the extent to which the workhorse DSGE model with financial frictions—the New Keynesian model of Christiano et al. [2005] (CEE hereafter) and Smets and Wouters [2007] (SW herafter) and augmented with the financial accelerator mechanism of Bernanke et al. [1999] (BGG hereafter)—can replicate the dynamics of the U.S. economy during the 2007–09 period.

We do so in two steps. First, following the recent work of Gilchrist and Zakrajšek [2010], we use micro-level prices of bonds issued by U.S. financial institutions to decompose financial credit spreads 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*. We show that the excess bond premium in the U.S. financial sector contains substantial predictive content for future economic activity, especially for the cyclically sensitive components of aggregate demand. Using an identified vector autoregression (VAR) framework, we demonstrate that shocks to the excess bond premium that are orthogonal to the current state of the economy, the net worth position of the nonfinancial sector, and the Treasury term structure cause economically and statistically significant declines in real economic activity.

In the second part of the paper, we use fluctuations in the estimated excess bond premium as a proxy for exogenous disturbances to the financial sector within the CEE/SW framework augmented with the BGG financial accelerator. We calibrate the key parameters of the model, so that the responses of macroeconomic aggregates to a financial shock match the corresponding impulse responses estimated using the actual data. Using this calibration, we then explore the extent to which observable fluctuations in the excess bond premium can account for macroeconomic dynamics during the recent financial crisis. Our results indicate that the model can fully account for the overall drop in consumption, investment, output, and hours worked that was observed during the crisis period. The model also does well at matching the observed decline in inflation and nominal interest rates, as well as the sharp widening of nonfinancial credit spreads.

Finally, we use this framework to analyze the potential benefits of an alternative monetary policy rule, a robust first-difference rule that allows for nominal interest rates to respond to changes in financial conditions as measured by the fluctuations in credit spreads. Our results suggest that by allowing the nominal interest rate to respond to credit spreads, as suggested recently by Taylor [2008] and McCulley and Toloui [2008], monetary policy can effectively dampen the negative consequences of financial market shocks on real economic activity, while experiencing only a modest increase in inflation.

The remainder of the paper is organized as follows. Section 2 contains a discussion of our data sources. Section 3 presents the empirical methodology used to estimate the excess bond premium and examines its predictive power for future economic activity. Section 4 outlines the general equilibrium framework used to study the impact of financial shocks on the macroeconomy and presents the corresponding results. Section 5 concludes.

2 Data Sources and Methods

2.1 Credit Spreads

The key information underlying our analysis comes from a sample of fixed income securities issued by U.S. financial corporations.[2] Specifically, for the period from January 1985 to June 2010, we extracted from the Lehman/Warga (LW) and Merrill Lynch (ML) databases month-end prices of outstanding financial corporate bonds that are actively traded in the secondary market.[3] To guarantee that we are measuring borrowing costs of different firms at the same point in their capital structure, we restricted our sample to senior unsecured issues with a fixed coupon schedule only.

We focus on the period from the mid-1980s onward, a period marked by a significant deregulation of financial markets (e.g., the repeal of Regulation Q (1986); the Riegle-Neal Act (1994); the Gramm-Leach-Bliley Act (1999)). In addition, rapid advances in information technology over the past quarter of the century have lowered the information and monitoring costs of investments in public securities, thereby increasing the tendency for corporate borrowing to take the form of negotiable securities issued directly in capital markets. By improving liquidity in both the primary and secondary markets, these changes in the financial landscape have facilitated more efficient price discovery and have likely improved the information content of credit spreads for future economic outcomes.[4]

The micro-level aspect of our data set allows us to construct credit spreads that are not biased by the maturity/duration mismatch, a problem that plagues credit spread indexes constructed with aggregated data. In particular, for each individual bond issue

[2] Our definition of the financial sector encompasses publicly-traded financial firms in the following 3-digit NAICS codes: 522 (Credit Intermediation & Related Activities); 523 (Securities, Commodity Contracts & Other Financial Investments & Related Activities); 524 (Insurance Carriers & Related Activities); and 525 (Funds, Trusts & Other Financial Vehicles). Government-sponsored entities, such as Fannie Mae and Freddie Mac, are excluded from the sample.

[3] These two data sources are used to construct benchmark corporate bond indexes used by market participants. Specifically, they contain secondary market prices for a vast majority of dollar-denominated bonds publicly issued in the U.S. corporate cash market. The ML database is a proprietary data source of daily bond prices that starts in 1997. By contrast, the LW database of month-end bond prices has a somewhat broader coverage and is available from 1973 through mid-1998 (see Warga [1991] for details).

[4] The ability of corporate bond credit spreads to predict economic activity has been documented by Gertler and Lown [1999]; King et al. [2007]; Mueller [2007]; Gilchrist et al. [2009b]; Gilchrist and Zakrajšek [2010]; and Faust et al. [2010].

in our sample, we construct a theoretical risk-free security that replicates exactly the promised cash-flows of the corresponding corporate debt instrument. For example, 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 in period t 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 a corresponding risk-free security—denoted by $P_t^f[k]$—we discount the promised cash-flow sequence $\{C(s) : s = 1, 2, \ldots, S\}$ using continuously-compounded zero-coupon Treasury yields in period t, obtained from the daily estimates of the U.S. Treasury yield curve reported 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 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 "duration mismatch" that would occur were the spreads computed simply by matching the corporate yield to the estimated yield of a zero-coupon Treasury security of the same maturity.

To ensure that our results are not driven by a small number of extreme observations, we eliminated all bond/month observations with credit spreads below 5 basis points and with spreads greater than 3,500 basis points. In addition, we dropped from our sample all observations with a remaining term-to-maturity of less than one year or more than 30 years, because calculating spreads for maturities of less than one year and more than 30 years would involve extrapolating the Treasury yield curve beyond its support. These selection criteria yielded a sample of 886 individual securities between January 1985 and June 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 193 firms.

Table 1 contains summary statistics for the key characteristics of bonds in our sample. Note that a typical financial firm may have 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 43 different 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 \$9.2 million to more than \$4.3 billion. Not surprisingly, the maturity of these debt instruments is fairly long, with the average maturity at issue of about 10 years. Because these securities tend to generate significant cash flow in the form of regular coupon payments, their average effective duration is only about 5.5 years.

According to the S&P's credit-rating scale, our sample spans a wide spectrum of credit quality, from "double CC" to "triple A." At "A2," however, the median bond is well within the investment-grade category, an indication of the generally high creditworthiness of financial firms, at least as perceived by one of the major rating agencies. Turning to returns, the (nominal) coupon rate on these bonds averaged 6.89 percent during our sample period, while the average expected total return, as measured by the nominal effective yield, was 6.78 percent per annum. Relative to Treasuries, an average bond in our sample has an expected return of about 172 basis points above the comparable risk-free rate, with the standard deviation of 253 basis points.

Figure 1 depicts the time-series evolution of credit spreads for our sample of bonds. With the exception of the recent financial crisis, the median credit spread on bonds issued by financial institutions—although countercyclical—fluctuated in a relatively narrow range. In spite of focusing on a relatively narrow segment of the U.S. financial system—namely, the publicly-traded financial corporations with senior unsecured debt trading in the secondary market—the interquartile range indicates a fair amount of dispersion in the price of debt across different institutions, information that is potentially useful for identifying shocks to the financial system.

2.2 Default Risk

We now turn to the construction of variables used to measure default risk in the financial sector, the crucial input in the construction of the excess bond premium. As in Gilchrist and Zakrajšek [2010], we employ the "distance-to-default" (DD) framework to measure a firm-specific probability of default at each point in time.[5] In this contingent claims approach to corporate credit risk—developed in the seminal work of Merton [1974]—it is assumed that a firm has just issued a single zero-coupon bond of face value D that matures at date T. Rational stockholders will default at date T only if the total value of the firm $V_T < D$; by assumption, the rights of the bondholders are activated only at the maturity date, as stockholders will maintain control of the firms even if the value of the firm $V_s < D$ for some $s < T$. In this context, the probability of default is, therefore, given by

[5] Several other papers consider similar market-based indicators of default risk for financial institutions. For example, Gropp et al. [2006] construct a distance-to-default for a sample of large European banks and find that the DDs are unbiased indicators of banks' financial health and have substantial predictive power for subsequent rating changes; Basurto et al. [2006] analyze the effect of financial and real variables on the probabilities of default—derived from the Merton DD-framework—for banks in the sample of OECD countries; Chan-Lau and Sy [2007] introduce a "distance-to-capital," a market-based measure of default risk for the commercial banking sector that accounts for fact banks are typically closed well before the equity has been completely wiped out; and Carlson et al. [2008] develop an indicator of financial distress based on the DDs for a sample of large U.S. financial institutions—including both commercial and investment banks—and find that the health of the financial sector has important implications for the real economy.

$\Pr[V_T < D]$, a quantity that depends on the value of the firm and is not directly observable.

The key insight of the contingent claims approach to 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. Thus, while the value of the firm cannot be directly observed, it 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. Specifically, it is assumed that the value of the firm follows a geometric Brownian motion:

$$dV = \mu_V V dt + \sigma_V V dW,$$

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. In that case, the Black-Scholes-Merton option-pricing framework implies that the value of the firm's equity satisfies:

$$E = V\Phi(\delta_1) - e^{-rT} D\Phi(\delta_2), \tag{1}$$

where r denotes the instantaneous risk-free interest rate, Φ is the cumulative standard normal distribution function, and

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

From equation (1) it is clear that 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:

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

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

$$\sigma_E = \left[\frac{V}{E}\right] \Phi(\delta_1)\sigma_V. \tag{2}$$

Using the observed values of E, D, σ_E, and r, equations (1) and (2) can be solved for V and σ_V using standard numerical techniques. However, as pointed out by Crosbie and Bohn [2003] and Vassalou and Xing [2004], the excessive volatility of market leverage (V/E) in equation (2) 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 recently proposed by Bharath and Shumway [2008]. Assuming a forecasting horizon of one year (i.e., $T = 1$) and letting the risk-free rate r equal the daily 1-year constant-maturity Treasury yield, we implement the model as follows: First, 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.[6] Second, we estimate σ_E from daily stock returns over the previous year and use this estimate to initialize $\sigma_V = \sigma_E[D/(E + D)]$. We then use this value of σ_V in equation (1) to infer the market value of the firm's assets V for every day of the previous year. Lastly, 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. The procedure is iterated on σ_V until convergence. The resulting solutions are then used to calculate the firm-specific DD at month-end as

$$DD = \frac{\ln(V/D) + (\mu_V - 0.5\sigma_V^2)}{\sigma_V}. \tag{3}$$

The corresponding implied probability of default is given by

$$\Pr[V \leq D] = \Phi(-DD) = \Phi\left(-\left(\frac{\ln(V/D) + (\mu_V - 0.5\sigma_V^2)}{\sigma_V}\right)\right), \tag{4}$$

which, under the assumptions of the Merton model, should be a sufficient statistic for predicting defaults.

We employ this methodology to calculate the distance-to-default for all U.S. financial corporations covered by the S&P's Compustat and CRSP (i.e., 2,477 firms over the Jan1985–June2010 period). Figure 2 plots the cross-sectional median of the DDs for the 192 bond issuers in our sample. As a point of comparison, the figure also depicts the cross-sectional median and interquartile range of the DDs for the entire Compustat-CRSP matched sample of financial firms.[7] According to this metric, the credit quality of the median bond issuer in our sample is, on average, comparable to that of the median financial firm. The median DD for both groups of firms is also strongly procyclical, implying—from the perspective of equity investors—a significant increase in the likelihood of default during economic downturns. Indeed, during the height of the recent financial crisis in late 2008 and early 2009, default risk in the financial sector reached a record level by recent historical standards.

The data shown in Figure 2 raise an obvious question: How good of an indicator of

[6]This assumption for the "default point" is also used by Moody's/KMV in the construction of their Expected Default Frequencies (EDFs) based on the Merton DD-model, and it 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.

[7]To ensure that our results were not driven by a small number of extreme observations, we eliminated from our sample all firm/month 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.

default risk is the distance-to-default, a measure that treats the firm's financial policy in such a highly simplified manner? This question is especially relevant in light of the fact that structural default models, such as the Merton DD-framework, do a rather poor job of pricing corporate bonds.[8] At first glance, the failure of structural models to explain the level of corporate credit spreads is likely due to their inability to predict accurately the likelihood of default. Indeed, much research on corporate credit risk over the past 25 years has been devoted to improving the way in which economists model default, including the introduction of stochastic default boundaries and dynamic capital structure (Leland and Toft [1996]), as well as accounting for the ability of the firm to alter its liability structure in the future (Anderson and Sundaresan [1996] and Collin-Dufresne and Goldstein [2001]). However, as shown by Huang and Huang [2003] and Eom et al. [2004], these important extensions of the contingent claims approach to corporate credit risk have so far failed to improve substantially the ability of such models to explain the level of corporate bond prices.

An alternative possibility is that factors unrelated to default risk—and, therefore, absent from the structural models altogether—play a significant role in the determination of bond prices. It is well known in the corporate finance literature 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] and Huang and Huang [2003]). Moreover, as shown by Collin-Dufresne et al. [2001], key variables from the structural model explain only a small portion of the variation in credit spreads, with the unexplained portion likely reflecting some combination of time-varying liquidity premium and fluctuations in the market price of default risk. Thus, it may be plausible that structural models might account well for the idiosyncratic default-risk component of corporate bond prices, while, at same time, fail to capture a significant portion of the total credit spread or the variability of bond returns.

Indeed, in a recent paper, Schaefer and Strebulaev [2008] present compelling micro-level evidence that even the simplest structural default model—the Merton's DD-model with nonstochastic interest rates—can account quite well for the default-risk component of corporate bond prices. In particular, they show that the standard structural 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

[8]In the corporate credit risk nomenclature, structural models of default employ contingent claims analysis to determine the appropriate price of a given risky bond; reduced-form models, in contrast, abstract from specifying the process for the value of the firm and model default as a pure jump process; see Lando [2004] for a recent review of the literature. By studying the empirical performance of a number of structural models, Eom et al. [2004] find a significant pricing error in both directions—some models tend to systematically overvalue, while other models tend to undervalue corporate debt.

credit risk, structural default models are able to capture the default-related component of credit spreads.[9] Thus in the context of pure default-risk exposure, it seems impossible to explain both the level of credit spreads and the sensitivity of corporate debt to equity—as argued by Schaefer and Strebulaev [2008], either the sensitivity to equity or the market price of default risk needs to be substantially higher to reconcile observed credit spreads with the contingent claims framework.

3 The Excess Bond Premium

In this section, we present a flexible empirical framework used to decompose financial credit spreads into two components: a default-risk component capturing the usual countercyclical movements in expected defaults; and a non-default-risk component that we argue represents the cyclical fluctuations in the relationship between default risk and credit spreads—the so-called excess bond premium. We then examine the extent to which movements in the excess bond premium are informative about subsequent economic growth. Lastly, we consider the macroeconomic implications of financial shocks, identified by orthogonalizing shocks to the excess bond premium in a standard VAR framework.

3.1 The Empirical Framework

As discussed in the previous section, the existing empirical evidence supports the view that the distance-to-default, our measure of institution-specific default risk, accounts well for the default-risk component of credit spreads, while not being able to explain a substantial non-default-risk component, the main focus of this paper. Nonetheless, our decomposition is complicated by the fact that a significant portion of bonds in our sample are callable (see Figure 3). As shown by Duffee [1998], if the firm's outstanding bonds are callable, 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.[10]

To deal with this issue, we utilize the micro-level aspect of our bond data to control

[9] The results of Schaefer and Strebulaev [2008] are consistent with those of Leland [2004], who finds that the Leland and Toft [1996] model is able to capture relatively well the term structure of default probabilities; they are also consistent with those of Huang and Huang [2003], who show that structural models calibrated with reasonable parameter values are capable of producing default probabilities that accord reasonably well with historical data.

[10] For example, as the general level of interest rates in the economy increases, the option to call becomes less valuable, which accentuates the price response of callable bonds relative to that of non-callable bonds. As a result, a rise in interest rates will, ceteris paribus, compress the credit spreads of callable bonds more than the credit spreads of their non-callable counterparts. In addition, prices of callable bonds are likely to be 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.

directly for the effects of the Treasury term structure and interest rate uncertainty on the credit spreads of callable bonds when constructing the excess bond premium. Specifically, we consider the following empirical bond-pricing model:

$$\ln S_{it}[k] = (1 + CALL_i[k]) \times (\beta_0 + \beta_1 DD_{it} + \beta_2 DD_{it}^2 + \gamma' X_{it}[k]) + \\ CALL_i[k] \times (\theta_1 LEV_t + \theta_2 SLP_t + \theta_3 CRV_t + \theta_4 VOL_t) + RTG_{it}[k] + IND_i[k] + \epsilon_{it}[k], \quad (5)$$

where $CALL_i[k]$ is an indicator variable that equals one if bond k (issued by firm i) is callable and zero otherwise; DD_{it} denotes the estimated year-ahead distance-to-default for firm i; and $\epsilon_{it}[k]$ is a "bond-pricing error."[11] In our framework, credit spreads of callable bonds are allowed to depend separately on the level (LEV_t), slope (SLP_t), and curvature(CRV_t) of the Treasury yield curve, the three factors that summarize the vast majority of the information in the Treasury term structure, according to Litterman and Scheinkman [1991] and Chen and Scott [1993].[12] The credit spreads of callable bonds are also influenced by the uncertainty regarding the path of long-term interest rates, as measured by the option-implied volatility on the 30-year Treasury bond futures (VOL_t).

We also allow for a nonlinear effect of default risk on credit spreads by including a quadratic term of DD_{it} in the bond-pricing regression, thereby accounting for the nonlinear relationship between credit spreads and leverage documented by Levin et al. [2004].[13] The vector $X_{it}[k]$, in contrast, controls for the bond-specific characteristics that could influence credit spreads through either term or liquidity premiums, including the bond's duration ($\ln DUR_{it}[k]$), the amount outstanding ($\ln PAR_{it}[k]$), and the bond's (fixed) coupon rate ($\ln CPN_i[k]$). The bond-pricing regression also includes credit rating fixed effects ($RTG_{it}[k]$), which capture the "soft information" regarding the firm's financial health, information that is complementary to our option-theoretic measures of default risk; see, for example, Löffler [2004, 2007]. The (3-digit NAICS) industry fixed effects ($IND_t[k]$) are included to control for any potential (time-invariant) differences in recovery rates across the different segments of the financial industry.

By averaging across bonds/firms at each point in time, we can define the predicted average credit spread as

$$\widehat{\bar{S}}_t = \frac{1}{N_t} \sum_i \sum_k \widehat{S}_{it}[k],$$

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

[12] 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].

[13] As a robustness check, we also considered higher-order polynomials of the distance-to-default, but the inclusion of cubic and quartic terms had virtually no effect on our results.

where $\widehat{S}_{it}[k]$ is the predicted credit spread on bond k from the bond-pricing model (5) and N_t is the number of bond/firm observations in month t. The excess bond premium in month t is then defined by the following linear decomposition:

$$EBP_t = \bar{S}_t - \widehat{\bar{S}}_t,$$

where $\bar{S}_t$ denotes the average credit spread in month t.

Table 2 contains OLS estimates—and associated standard errors—of the key parameters from the bond-pricing model (5). According to the entries in the table, our option-theoretic measure of default risk is a highly significant determinant—both economically and statistically—of credit spreads on financial corporate bonds. Translating the coefficients on the distance-to-default into the impact of variation in default risk (the sum of the linear and quadratic *DD* terms) on the *level* of credit spreads, our estimates imply that an increase in the distance-to-default of one standard deviation—a sign of improving credit quality—leads to a narrowing of spreads of about 19 basis points for both callable and non-callable bonds.

The estimates in Table 2 also indicate that the shape of the Treasury term structure and interest rate uncertainty have first-order effects on the credit spreads of callable bonds, which are consistent with the theoretical predictions. For example, a one standard deviation increase in the level factor implies a 78 basis points reduction in the credit spreads on callable bonds, while a one standard deviation increase in the slope factor lowers credit spreads on such bonds 30 basis points. An increase in the option-implied volatility on the long-term Treasury bond futures of one percentage point implies a widening of callable credit spreads of 13 basis points, because the rise in interest rate uncertainty lowers the prices of callable bonds by boosting the value of the embedded call options.

The importance for controlling for the optionality of bonds is illustrated in Figure 4, which shows the time path of the average credit spread in our data set, along with the predicted values from the bond-pricing regression (5) and from the specification that imposes the restriction $\theta_1 = \ldots = \theta_4 = 0$, that is, a specification that does not control for the effects of the Treasury term structure and interest rate uncertainty. The former, clearly, fits the data much better. Fluctuations in the value of embedded call options had an especially significant effect during the recent financial crisis. The plunge in the risk-free interest rates and the steepening of the Treasury term structure that began with the onset of the financial crisis in the summer of 2007—two factors that more than offset the spike in (long-term) interest rate volatility that occurred during that period—imply higher predicted values for the credit spreads of callable bonds. As result, the option-adjustment terms account for more than 400 basis points of the total increase in the average credit spread during the height of the financial crisis in the autumn of 2008. In sum, as evidenced by both the adjusted R^2 in Table 2 and the predicted values in Figure 4, our empirical bond-pricing model explains a substantial portion of the cyclical fluctuations in financial credit spreads.

Figure 5 shows the estimated excess bond premium in the U.S. financial sector, the difference between the average financial credit spread and its predicted value from the bond-pricing regression (5). While clearly countercyclical, the excess bond premium appears to be a particularly timely indicator of strains in the financial system. The sharp run-up in the premium during the early 1990s, for example, is consistent with the view that capital pressures on commercial banks in the wake of the Basel I capital requirements significantly exacerbated the 1990–91 economic downturn by reducing the supply of bank credit (Bernanke and Lown [1991]). In contrast, the robust health of the financial system at the start of the 2001 recession has been cited as an important factor for the absence of a "credit crunch," which, in turn, likely contributed to the fact that the downturn remained localized in certain troubled industries, particularly the high-tech sector (Stiroh and Metli [2003]).

In regard to the recent financial crisis, the intensifying downturn in the housing market and the emergence of significant strains in the term funding markets in the United States in Europe during the summer of 2007 precipitated a sharp increase in the excess bond premium. At that time, banking institutions, in addition to their mounting concerns about actual and potential credit losses, recognized that they might need to take a large volume of assets onto their balance sheets, given their existing commitments to customers and the heightened reluctance of investors to purchasing an increasing number of securitized products. The recognition that the ongoing turmoil in financial markets could lead to substantially larger-than-anticipated calls on their funding capacity and investors' concerns about valuation practices for opaque assets were the primary factors behind the steady climb of the excess bond premium during the remainder of 2007 and over the subsequent year.

The full-fledged global nature of the crisis became apparent in the early autumn of 2008, when Lehman Brothers—with its borrowing capacity severely curtailed by a lack of collateral—filed for bankruptcy. Investor anxiety about financial institutions escalated sharply, and market participants became extraordinarily skittish and pulled back from risk-taking even further. Amid cascading effects of these financial shocks, which included a run on money market mutual funds, the government's rescue of AIG, and the failure of Washington Mutual, a large thrift, the excess bond premium shot up, reaching 325 basis points by October 2008.

Responding to the panic that was rapidly engulfing the entire global financial system, the Federal Reserve, at times acting in concert with foreign central banks, used its emergency and lending authorities to expand its existing liquidity facilities, while also announcing several additional initiatives aimed at restoring the confidence in the financial system.[14] Concomitantly, the Treasury announced a temporary guarantee program for money market mutual funds and proposed the Troubled Asset Relief Program (TARP), under which government funds were to be used to help stabilize the banking system.[15] Lastly, the Federal Deposit Insurance Corporation (FDIC) provided a temporary guarantee for selected senior unsecured obligations of participating depository institutions and many of their parent holding companies, as well as for all balances in non-interest-bearing transaction deposit accounts at participating depository institutions—the so-called Temporary Liquidity Guarantee Program (TLGP).

After unprecedented actions taken by the U.S. government during the September–October period and announcements of similar programs in a number of other countries, the panic abated and stresses in financial markets eased, as evidenced by the drop in the excess bond premium. Despite these improvements, investors remained greatly concerned about the soundness of financial institutions. As a result, pressures on already-strained balance sheets of financial institutions remained substantial and continued to threaten their viability, a situation that greatly impinged the flow of credit to businesses and households.[16] At the same time, the severity and persistence of financial strains and the significant tightening of aggregate credit conditions caused the downturn in economic activity that has been unfolding since late 2007 to accelerate noticeably. Reflecting these adverse macroeconomic developments, coupled with reports of large losses in the fourth quarter of 2008, the pressure on financial firms intensified during the first few months of 2009, sending the excess bond premium to a new record—330 basis points—by March 2009.

The monetary authorities again responded forcefully to these adverse financial developments: The FOMC kept its target for the federal funds rate between 0 and 1/4 percent, expanded direct purchases of agency debt and mortgage-backed securities (MBS), and began direct purchases of longer-term Treasury securities. Partly as a result of these actions, conditions in financial markets, although they remained strained, began to show signs of improvements in the late spring of 2009. Substantially better-than-expected first-quarter earnings results for some large financial institutions contributed importantly to improved in-

[14]The new initiatives included the creation of the Asset-Backed Commercial Paper Money Market Mutual Fund Liquidity Facility (AMLF), the creation of the Commercial Paper Funding Facility (CPFF), and the creation of the Money Market Investor Funding facility (MMIFF). A detailed description of all the new policy tools used by the Federal Reserve to address the recent financial crisis can be found at Credit and Liquidity Programs and the Balance Sheet.

[15]On October 3, 2008, the Congress approved and provided funding for TARP as part of the Emergency Economic Stabilization Act. Using funds from TARP, the Treasury established a voluntary capital purchase plan, under which the U.S. government was able to inject equity—in the form of preferred shares—into the banking system.

[16]In November 2008, Citigroup found itself under significant financial pressure. In response, the FDIC, the Treasury, and the Federal Reserve provided a package of loans and guarantees to bolster Citigroup's financial positions; a similar package was arranged for Bank of America in January.

vestor sentiment, and share prices of banks and insurance companies moved higher and their credit default swap (CDS) premiums declined. The release of the findings of the Supervisory Capital Assessment Program (i.e., the "stress test") in May further reduced uncertainty and restored confidence in the financial system, and banking organizations were able to issue significant amounts of equity and non-guaranteed debt during subsequent months.

The link between the excess bond premium and the health of the financial sector is made more explicit by Figure 6, which plots the premium against the (annualized) return on assets (ROA) in the U.S. financial corporate sector, calculated using the Compustat data. The high degree of negative comovement between the two series (correlation of -0.64) is consistent with the view that risk premiums on assets fluctuate closely in response to movements in capital and balance sheet conditions of financial intermediaries, a fact emphasized, documented, and analyzed extensively by Adrian and Shin [2010] and Adrian et al. [2010a,b].

From a theoretical perspective, recent work by He and Krishnamurthy [2009, 2010] shows that adverse macroeconomic conditions, by depressing the capital base of financial intermediaries, can reduce the risk-bearing capacity of the marginal investor, causing a sharp increase in the conditional volatility and correlation of asset prices and a drop in risk-free interest rates. Relatedly, Acharya and Viswanathan [2007] develop a framework in which financial intermediaries—in response to a sufficiently severe aggregate shock—are forced to de-lever by selling their risky assets to better-capitalized firms, causing asset markets to clear only at "cash-in-the-market" prices (cf. Allen and Gale [1994, 1998]). The recent work by Brunnermeier and Pedersen [2009] and Garleanu and Pedersen [2009], in contrast, explores how margins or haircuts—the difference between the security's price and collateral value that must be financed with the trader's own capital—interact with liquidity shocks in determining the dynamics of asset prices.

In light of the above discussion, we now examine whether fluctuations in other commonly-used indicators of financial market stress are informative about the subsequent movements in the excess bond premium. Specifically, we regress the excess bond premium on its own lagged value and a lagged value of one of the following asset market indicators: (1) the option-implied volatility on the S&P 100 stock price futures; (2) the implied volatility on the 3-month Eurodollar futures; (3) the implied volatility on the 30-year Treasury bond futures; (4) the difference between the 3-month Libor and the 3-month Treasury bill rate (the TED spread); (5) the difference between the 5-year swap rate and the 5-year Treasury yield; and (6) the yield spread between the off-the-run and on-the-run 10-year Treasury notes (the off/on-the-run spread).

According to the results reported in Table 3, measures of implied volatility in both the equity and fixed income markets—proxies for time-varying economic uncertainty—have no predictive power for the excess bond premium.[17] Fluctuations in the TED spread—the conventional measure of counterparty credit risk in the interbank funding markets—also appear to be unrelated to the subsequent movements in the excess bond premium. The direct effects of counterparty risks in the Libor market, however, are likely to be small, reflecting both the marking to market convention and collateralization requirements, as well as the generally high credit quality of market participants. As emphasized by Joslin and Singleton [2009], investor concerns about the possible deterioration in the credit quality of financial intermediaries, especially those that insure bonds, often manifest themselves through higher risk premiums in the interest rate swap market. Indeed, the 5-year swap-Treasury spread has some marginal predictive power for the excess bond premium, although the improvement in the goodness-of-fit is rather modest ($\bar{R}^2 = 0.652$ vs. $\bar{R}^2 = 0.618$ from an AR(1) model).[18] Finally, the off/on-the-run Treasury spread—a gauge of investor liquidity preference (Krishnamurthy [2002])—is also uninformative about the future movements in the excess bond premium.

3.2 The Excess Bond Premium and Macroeconomic Dynamics

We now turn to the predictive power of the excess bond premium for future economic activity. Defining

$$\Delta^h Y_{t+h} \equiv \frac{400}{h} \ln\left(\frac{Y_{t+h}}{Y_t}\right),$$

where Y_t denote a measure of economic activity in quarter t and h the forecast horizon, we estimate the following forecasting regression:

$$\Delta^h Y_{t+h} = \alpha + \sum_{i=0}^{p} \beta_i \Delta Y_{t-i} + \gamma_1 EBP_t + \gamma_2 TS_t + \gamma_3 RFF_t + \epsilon_{t+h}. \tag{6}$$

Thus when analyzing the predictive content of the excess bond premium (EBP) for economic activity, we control for the current stance of monetary policy as measured by the "term spread" (TS)—that is, the slope of the Treasury yield curve, defined as the difference between the 3-month and 10-year constant-maturity Treasury yields—and the real federal funds rate (RFF).[19]

[17] Recent work by Gilchrist et al. [2010] examines, both empirically and theoretically, the macroeconomic effects of fluctuations in economic uncertainty in the presence of financial market frictions.

[18] The swap-Treasury spread is a rough measure of the swap risk premium. As shown by Krishnamurthy and Vissing-Jorgensen [2010] private yield spreads measured relative to comparable-maturity Treasuries can be influenced importantly by the relative supply and demand pressures in the Treasury market. In addition, Treasury yields embody a substantial convenience premium. These factors, however, do not influence the level of swap yields. As a result, Joslin and Singleton [2009] advocate estimating swap risk premiums using a dynamic term structure model proposed by Joslin et al. [2009].

[19] The term spread is used often as an indicator of the stance of monetary policy—the higher the term spread, the more restrictive is the current stance of monetary policy and, hence, the more likely is economy

We estimate the forecasting regression (6) by OLS, with the lag length p determined by the Akaike Information Criterion (AIC).[20] Table 4 details the predictive power of the excess bond premium for the growth rate of real GDP at both the 1- and 4-quarter ahead forecast horizons. Entries in the table correspond to standardized estimates—and associated t-statistics—of the coefficients associated with the financial indicators as well as the in-sample goodness-of-fit as measured by the adjusted R^2.

In the very near term, neither indicator of the stance of monetary policy is informative about economic growth prospects, a finding that is not too surprising given the fact that monetary policy affects the real economy with a lag. The excess bond premium, in contrast, contains substantial predictive power for near-term economic growth. The point estimate of -0.500 implies that an increase in the excess bond premium of 1 standard deviation—about 50 basis points—leads to a reduction in the (annualized) growth of real GDP of almost 1.25 percentage points in the subsequent quarter. At the year-ahead forecast horizon, the Treasury term structure has significant predictive ability for output growth, with a flat or inverted yield curve signalling a slowdown in economic activity. The excess bond premium, however, remains—both statistically and economically—a highly significant predictor of future economic growth. According to our estimates, a 1 standard deviation increase in the excess bond premium in quarter t implies a deceleration in real GDP of almost a full percentage point over the subsequent four quarters.

The results reported in Table 5 examine the predictive ability of the excess bond premium for the main categories of personal consumption expenditures and private investment. At the 1-quarter horizon (top panel), movements in the excess bond premium are highly informative about the growth in all major components of business investment as well as for the growth of personal consumption expenditures on both durable and nondurable goods, where the latter includes spending on services. The slope of the yield curve, in contrast, is somewhat informative about the near-term swings in residential investment, the one component of private domestic final purchases for which the excess bond premium has no

to decelerate in subsequent quarters. In general, however, the shape of the yield curve contains information about term premiums and the average of expected future short-term interest rates over a relatively long horizon. As emphasized by Hamilton and Kim [2002] and Ang et al. [2006], the term premium and expectations hypothesis components of the term spread have very different correlations with future economic activity. The real federal funds rate, in contrast, is a measure of the stance of monetary policy that is relatively unadulterated by the effects of time-varying term premiums. In calculating the real funds rate, we employ a simplifying assumption that the expected inflation is equal to lagged core PCE inflation. Specifically, the real funds rate in quarter t is defined as the average effective federal funds rate during that quarter less realized inflation, where realized inflation is given by the log-difference between the core PCE price index in quarter $t-1$ and its lagged value a year earlier.

[20] For the forecasting horizons $h > 1$, the MA$(h-1)$ structure of the error term ϵ_{t+h} induced by overlapping observations is taken into account by computing the covariance matrix of regression coefficients according to Hodrick [1992]. In the case of non-overlapping data (i.e., $h = 1$), our inference is based on the heteroscedasticity-consistent asymptotic covariance matrix (HC3) computed according to MacKinnon and White [1985].

predictive power. Although the forecasting ability of the shape of the Treasury term structure improves noticeably at the year-ahead forecast horizon (bottom panel), the excess bond premium remains, statistically and economically, a highly significant predictor of economic activity, especially in the business sector. Indeed, for some of the most cyclically volatile macroeconomic variables, such as inventory investment and E&S spending, the economic impact of the excess bond premium is about thrice as large as that of the term spread.

To examine macroeconomic consequences of financial shocks, we include the excess bond premium into an otherwise standard VAR. Our specification includes the following endogenous variables: (1) consumption growth as measured by the log-difference of real personal consumption expenditures on nondurable goods and services; (2) investment growth as measured by the log-difference of real private investment (residential and business) in fixed assets; (3) output growth as measured by the log-difference of real GDP; (4) the log-difference in hours worked; (5) inflation as measured by the log-difference of the GDP price deflator; (6) the growth of the market value of net worth in the nonfinancial (nonfarm) corporate sector; (7) the 10-year (nominal) Treasury yield; (8) the effective (nominal) federal funds rate; and (9) the excess bond premium.[21]

This multivariate framework allows to trace out the effect of a shock to the excess bond premium that is orthogonal to measures of economic activity and inflation, the balance sheet position of the nonfinancial sector, and the level of short- and long-term interest rates. The dynamic responses of the key macroeconomic aggregates to an impact of such a financial shock also serve as useful benchmark for the calibration of the DSGE model, considered in the next section. We estimate the VAR over the 1985:Q1–2010:Q2 period, using two lags of each endogenous variable.

Figure 7 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—almost 30 basis points—causes a significant slowdown in economic activity. In economic terms, the implications of this adverse financial shock are substantial: Although the decline in consumption is relatively mild, total private fixed investment drops significantly, bottoming out a full percentage point below trend about five quarters after the shock; hours worked also decelerate markedly, and the output of the economy as a whole does not begin to recover until about a year and a half after the initial impact.

The downturn in economic activity is amplified in part by the substantial drop in the net worth of nonfinancial firms, and the repair of corporate balance sheets is slow and protracted. The combination of the economic slack and appreciable disinflation in the wake

[21] Consumption and investment series are constructed from the underlying NIPA data using the chain-aggregation methods outlined in Whelan [2002]. The market value of net worth is taken from the U.S. Flow of Funds Accounts.

of the financial shock elicits a significant easing of monetary policy, as evidenced by the decline in the federal funds rate. As shown in Figure 8, these financial shocks also account for about 10 percent of the variation in economic activity at business cycle frequencies, a non-trivial proportion and one that is comparable to the amount of variation typically attributed monetary shocks.

4 General Equilibrium Model

In this section, we describe a New Keynesian general equilibrium model that allows for financial frictions along the lines of the financial accelerator mechanism formulated by Bernanke et al. [1999]. In absence of financial frictions, the model reduces to a standard New Keynesian framework of the type developed by Christiano et al. [2005] and analyzed subsequently by Smets and Wouters [2007] and Justiniano et al. [2010]. Key features of the CEE/SW framework include habit formation, higher-order adjustment costs to investment, fixed-costs of production and variable capacity utilization, and nominal rigidities due to a Calvo price-setting mechanism with indexation. Monetary policy is conducted via a Taylor-type rule that sets the nominal interest rate as a function of inflation, output growth, and lagged nominal rates.

The estimated versions of the CEE/SW framework, augmented with the BGG financial accelerator mechanism, can be found in the recent work by Christiano et al. [2009] and Gilchrist et al. [2009a]. These papers show that allowing for *unobservable* shocks to the financial sector can account for a substantial fraction of the variability in investment and output in U.S. historical data. In this paper, by contrast, we use fluctuations in the estimated excess bond premium as a proxy for exogenous disturbances to the financial sector, which boost the cost of external finance for nonfinancial borrowers.

In addition to augmenting the CEE/SW framework with the BGG financial accelerator mechanism, we also modify the preferences of the representative household to better match the dynamics of consumption over the course of the business cycle. Specifically, our modification nests standard household preferences—in which consumption and leisure are separable—with those proposed by Greenwood et al. [1988] (GHH hereafter), a specification in which some fraction of labor disutility is quasi-linear in consumption and is, therefore, embedded in the household habit. It is well known that with preferences that are separable between consumption and leisure, financial shocks produce negative comovement between consumption and investment, a counterfactual response that is only partially mitigated by the model's New Keynesian features. With GHH preferences and habit, by contrast, financial shocks lead to positive comovement between consumption and investment.

We choose the key parameters of the model, so that the responses of macroeconomic aggregates to a financial shock match the corresponding impulse responses shown in Figure 7. Using this calibration, we then explore the extent to which observable fluctuations in the excess bond premium can account for macroeconomic dynamics during the recent financial crisis. By choosing parameters to match the estimated impulse response functions, our results imply that the model can fully account for the overall drop in consumption, investment, hours, and output that was observed during the crisis period. The model also does well at matching the observed decline in inflation and nominal interest rates, as well as the sharp widening of nonfinancial credit spreads. We then use this framework to analyze the potential benefits of an alternative monetary policy rule, a rule that allows for nominal interest rates to respond to changes in financial conditions as measured by the movements in credit spreads.

4.1 The Agents

Households: The representative household maximizes the present discounted value of per-period utility:

$$E_t\left\{\sum_{s=0}^{\infty}\beta^s U(Z_t, Z_{t-1}, L_t)\right\},$$

where L_t denotes labor supply, Z_t is the habit index, and preferences are assumed to satisfy:

$$U(Z_t, Z_{t-1}, L_t) = \ln(Z_t - \gamma Z_{t-1}) + (1-\omega)v(L_t),$$

with

$$\begin{aligned} Z_t &= C_t - \omega v(L_t); \quad 0 \leq \omega \leq 1, \\ v(L) &= \frac{\kappa}{1+\eta}L^{1+\eta}. \end{aligned}$$

These preferences allow for internal habit formation in a manner that nests standard preferences—in which labor is separable from consumption and, therefore, the habit index Z_t depends only on past consumption (i.e., $\omega = 0$)—with GHH preferences in which the index Z_t is a quasi-linear combination of consumption and labor; note that in the latter case, the labor supply decision is independent of the wealth effect, regardless of the degree of habit formation.

Households maximize their objective subject to an inter-temporal budget constraint:

$$C_t + b_t\frac{B_{t+1}}{P_t} = W_t L_t + \frac{B_t}{P_t} + DIV_t - T_t,$$

where P_t denotes the price level, b_t is the discount price of nominal bonds, W_t is the real wage, DIV_t denotes dividends earned from imperfectly competitive retail firms, and T_t are the lump-sum taxes used to finance government spending and net transfers.

Letting

$$\lambda_t = \frac{1}{Z_t - \gamma Z_{t-1}} \quad \text{and} \quad R^n_{t+1} = \frac{1}{b_t}$$

denote the marginal utility of wealth and the gross nominal interest rate, respectively, household optimality conditions then imply the following inter-temporal savings condition:

$$\lambda_t = E_t \left\{ \beta \lambda_{t+1} R^n_{t+1} \frac{P_t}{P_{t+1}} \right\};$$

and the following intra-temporal labor supply condition:

$$\lambda_t \left[W_t - \omega v'(L_t) \right] = (1-\omega) v'(L_t).$$

When $\omega = 0$, this reduces to the standard labor supply condition:

$$\lambda_t W_t = v'(L_t),$$

whereas when $\omega = 1$, we obtain the GHH specification, in which labor supply is independent of the marginal utility of wealth:

$$W_t = v'(L_t).$$

Retail firms: We assume the existence of a continuum of retail firms—indexed by τ—that supply, at a monopolistic price P^τ_t, a retail good Y^τ_t, produced using the homogeneous wholesale output Y^w_t that is purchased in a competitive market at price P^w_t. All retailers face a fixed cost of production Ξ and have access to a technology that allows them to transform wholesale goods into retail goods in a one-for-one manner. The final-good output is a Dixit-Stiglitz aggregate of the differentiated retail goods:

$$Y_t = \left(\int_0^1 Y_t^{\tau\,\frac{\varepsilon-1}{\varepsilon}} d\tau \right)^{\varepsilon/(\varepsilon-1)};$$

and the price level is given by

$$P_t = \left(\int_0^1 P_t^{\tau(1-\varepsilon)} d\tau \right)^{1/(1-\varepsilon)};$$

the individual retail demand satisfies

$$Y^\tau_t = \left(\frac{P^\tau_t}{P_t} \right)^{-\varepsilon} P_t;$$

where ε determines the elasticity of demand.

Retail firms face nominal price rigidities that evolve according to a standard Calvo price-setting process, in which prices can be adjusted with constant probability $\theta > 0$ in each period. In periods when prices are not adjusted, we assume that the nominal price is fully indexed to past inflation. The price reset in period t—denoted by P_t^*—is given by

$$P_t^* = \left[\frac{\varepsilon}{\varepsilon - 1}\right] \frac{E_t \sum_{i=0}^{\infty} \theta^i \beta^i \left(\frac{\lambda_{t+i}}{\lambda_t}\right)^{-1} P_t^w Y_{t+i} \left(\frac{1}{P_{t+i}}\right)^{\varepsilon}}{E_t \sum_{i=0}^{\infty} \theta^i \beta^i \left(\frac{\lambda_{t+i}}{\lambda_t}\right)^{-1} Y_{t+i} \left(\frac{1}{P_{t+i}}\right)^{\varepsilon}}.$$

The price level in period t depends on both the price level in the previous period indexed to lagged inflation and on the newly set price P_t^*:

$$P_t = \left[\theta \left((1 + \pi_{t-1}) P_{t-1}\right)^{1-\varepsilon} + (1 - \theta)(P_t^*)^{1-\varepsilon}\right]^{\frac{1}{1-\varepsilon}}.$$

Entrepreneurs: As in BGG, entrepreneurs are long-lived, risk-neutral agents that purchase capital and produce wholesale goods that are sold to retail firms. Entrepreneurs purchase capital K_t at time $t-1$ in a competitive market at the relative price Q_{t-1} and make labor and capital utilization decisions at time t to maximize profits. Purchases of capital goods are financed using a combination of entrepreneurial net worth and debt issued to the household sector. The presence of financial frictions in the process of issuing debt implies that entrepreneurs must pay a premium on external funds raised from households. This premium is an increasing function of the leverage of the entrepreneurial sector. Because entrepreneurs are long-lived, the net worth of the entrepreneurial sector is a state variable of the economy.

Entrepreneurs are also assumed to "die" with constant exogenous probability in each period, in which case, they are replaced within the period by new entrepreneurs. As a result, entrepreneurs discount the future more heavily than households and withhold from consuming until they die. These assumptions imply that entrepreneurs will not, on average, accumulate enough savings to fully finance their investment opportunities. We assume that upon death, entrepreneurial consumption is fully taxed and rebated lump-sum to the household sector. New entrepreneurs are given (negligible) start-up funds to begin operation; these funds are assumed to be provided via a lump-sum tax on households.

Entrepreneurs choose labor L_t and capital utilization u_t to maximize profits

$$\left(\frac{P_t^w}{P_t}\right) Y_t^w - W_t L_t - \psi(u_t) K_t,$$

subject to the Cobb-Douglas production function:

$$Y_t^w = A_t \left(K_t^s\right)^\alpha L_t^{1-\alpha},$$

where $K_t^s = u_t K_t$ denotes capital services and $\psi(u_t)K_t$ is the resource cost of increasing the rate of capital utilization. Letting P_t^k denote the entrepreneur's marginal profitability of capital, then cost minimization implies

$$\frac{W_t L_t}{P_t^k K_t^s} = \frac{1-\alpha}{\alpha},$$

where

$$P_t^k = \psi'(u_t).$$

The entrepreneur's marginal cost of production is given by:

$$MC_t = \frac{P_t^W}{P_t} = \frac{1}{A_t} W_t^{1-\alpha} \left(P_t^k\right)^\alpha \left(\alpha^{-\alpha}(1-\alpha)^{-(1-\alpha)}\right).$$

Because capital goods are resold at time $t+1$, the real rate of return on capital—denoted by R_{t+1}^k—is the sum of the marginal profitability of capital and the capital gain:

$$R_{t+1}^k = \frac{\left[u_{t+1} P_{t+1}^k - \psi(u_{t+1}) + (1-\delta) Q_{t+1}\right]}{Q_t},$$

$0 < \delta < 1$ is the rate of capital depreciation.

4.2 Closing the Model

The external finance premium: The external finance premium—denoted by s_t—is defined as the ratio of the expected real rate of return on capital (which, in equilibrium, is equal to the cost of external funds) to the expected real rate of return on a riskless bond, which is interpreted as the cost of internal funds:

$$s_t \equiv \frac{E_t R_{t+1}^k}{E_t \left[R_{t+1}^n \frac{P_t}{P_{t+1}}\right]}.$$

In order to focus on the primary distortion associated with financial frictions—namely, the introduction of a time-varying countercyclical wedge between the rate of return on capital and the rate of return on the riskless bond held by households—we adopt a number of simplifications with respect to the original BGG formulation. Specifically, financial market imperfections imply that the external finance premium increases when the leverage of the borrowers increases:

$$s_t = \exp(\sigma_t)\left(\frac{Q_t K_{t+1}}{N_{t+1}}\right)^{\chi}, \tag{7}$$

where $0 < \chi < 1$ is the parameter governing the strength of this response and σ_t denotes a shock to the financial intermediation process that is assumed to evolve according to

$$\sigma_t = (1-\rho_\sigma)\sigma + \rho_\sigma \sigma_{t-1} + \epsilon_t.$$

In our context, an increase in σ_t raises the cost of external finance for a given amount of leverage and hence may be interpreted as an increase in the effective cost of financial intermediation.[22]

Net worth at the beginning of period $t+1$ is the return on capital less the repayment of the loan made at the beginning of period t:

$$N_{t+1} = \mu\left[R_t^k Q_{t-1} K_t - s_{t-1} E_{t-1}\left[R_t^n \frac{P_{t-1}}{P_t}\right](Q_{t-1}K_t - N_t)\right] + (1-\mu)W^E.$$

where $0 < \mu < 1$ is the parameter reflecting the depletion of net worth owing to the death rate of entrepreneurs; W^E is the lump-sum transfer from households to new entrepreneurs; and $s_{t-1}E_{t-1}\left[R_t^n \frac{P_{t-1}}{P_t}\right]$ is the equilibrium (gross) interest rate on risky debt.

Capital goods production: We assume the existence of a competitive capital-goods producing sector that produces new capital goods according to an adjustment technology that is increasing in the rate of investment. Aggregate capital accumulation evolves according to

$$K_{t+1} = (1-\delta)K_t + \left[1 - S\left(\frac{I_t}{I_{t-1}}\right)\right] I_t,$$

where $S(\cdot)$ is the capital adjustment cost function. The optimality condition for capital goods producers implies the following relationship between the price of capital and the investment rate:

$$Q_t\left[1 - S\left(\frac{I_t}{I_{t-1}}\right)\right] = Q_t S'\left(\frac{I_t}{I_{t-1}}\right)\frac{I_t}{I_{t-1}} - E_t\left[\beta\frac{\lambda_{t+1}}{\lambda_t} Q_{t+1} S'\left(\frac{I_{t+1}}{I_t}\right)\frac{I_{t+1}}{I_t}\right].$$

Aggregate resource constraint: Let G denote government spending, which is financed through lump-sum taxes of households. The resource constraint implies that final output

[22] Christiano et al. [2009] interpret such a shock as an increase in the idiosyncratic variance of the firm-level project returns in the entrepreneurial sector. Such a shock raises the cost of external finance because in the Costly-State Verification (CSV) framework adopted by BGG, an increase in idiosyncratic uncertainty implies an increase in the cost of debt finance; see also Gilchrist et al. [2010]. Such a shock may also be interpreted as an increase in the cost of default in the CSV framework.

Y_t satisfies

$$Y_t = C_t + I_t + \psi(u_t)K_t + G + \Xi.$$

Monetary policy: For our baseline case, we assume that the policy interest rate depends on inflation and output growth. In letting the short-term interest rate respond to output growth—as opposed to the output gap—the monetary authority is assumed to follow a robust first-difference rule of the type proposed by Orphanides [2003]. Moreover, as shown by Orphanides and Williams [2006], such first-difference rules are highly successful in stabilizing economic activity in the presence of imperfect information regarding the structure of the economy.[23] As an alternative, we also consider a policy rule in which monetary authorities also respond to movements in the external finance premium according to

$$\ln R^n_{t+1} = \ln R^n + \phi_\pi \ln \pi_t + \phi_{\Delta y} \Delta \ln Y_t + \phi_s \ln s_t,$$

where $\phi_\pi > 0$, $\phi_{\Delta y} > 0$, and $\phi_s < 0$ are the parameters governing the sensitivity of the policy interest rate to inflation, output growth, and the external finance premium, respectively.

4.3 Calibration

We take the following approach to model calibration. Most parameters, including those that determine the model's steady-state (i.e., the household's labor supply elasticity and the degree of habit formation), are fixed at values commonly used in the literature. Parameters that govern the behavior of monetary policy and the financial shock process are directly estimated from the data. We then choose a subset of four key elasticities that determine the dynamics of inflation, investment, consumption, and output in order to match the model's impulse response function to those estimated using our VAR framework.

A period in the model is a quarter. We set the discount factor $\beta = 0.984$; the labor share of income $1 - \alpha = 2/3$; the labor supply elasticity is $1/v = 3$ and the degree of habit formation $\gamma = 0.75$. The depreciation rate $\delta = 0.025$. The steady-state markup is $\varepsilon/(\varepsilon - 1) = 1.1$. Government spending, as a share of output, is set to 0.2 in steady state. The parameters of the monetary policy rule are chosen to match the parameters from a regression of the nominal interest rate on its own lag, inflation, and output growth, estimated over the period 1985:Q1–2010:Q2. This regression yields $\rho_i = 0.92$, $\phi_\pi = 0.15$, and $\phi_{\Delta y} = 0.11$. Similarly, we choose the autoregressive coefficient for the financial shock to match an AR(1) regression coefficient of the excess bond premium over our sample sample period—this implies $\rho_\sigma = 0.75$ for the persistence of our shock process.

[23] According to the simulations reported by Orphanides and Williams [2006], such a robust monetary policy rule yields outcomes for the federal funds rate that are very close to those seen in the actual data, especially for the period since the mid-1980s.

The log-linearized model implies that there are two key financial parameters to choose—the steady-state leverage ratio and the elasticity of the external finance premium with respect to leverage. The steady-state ratio of the real value of the capital stock to the entrepreneur's net worth is chosen so that the steady-state leverage ratio is 100 percent, or $(QK - N)/N = 1$, which implies $K/N = 2$ (note that $Q = 1$ in steady state). There are four remaining parameters to calibrate: ω the share of labor disutility that enters the habit index Z_t; $\Lambda_I \equiv -S''(I_t/I_{t-1})/S'(I_t/I_{t-1})$, the elasticity of asset prices with respect to the investment-capital ratio; θ, the parameter determining the frequency of price resets; and χ, the elasticity of the external finance premium with respect to leverage. Increasing ω results in a greater comovement between consumption, employment, and output in response to a financial shock, while increasing Λ_I strengthens the hump-shaped response of investment to financial shocks. The size of χ governs the strength of the financial accelerator mechanism, with bigger values of χ leading to larger declines in investment and output in response to a financial shock.[24]

We choose these four parameters so that the model's impulse responses to a financial shock match the size of the decline in consumption, investment, output, and inflation in response to a shock to the excess bond premium documented in the previous section. This implies the following values for our four key elasticities: $\omega = 0.4$; $\Lambda_I = 3.0$, $\theta = 0.8$; and $\chi = 0.1$. The investment elasticity Λ_I and the Calvo price setting parameter θ are within the range of standard estimates in the literature. In contrast, the choice of $\chi = 0.1$, roughly twice as high as the calibration adopted by BGG, is somewhat higher than that suggested by the recent estimates. Although there are no prior estimates of ω, we note that if $\omega = 1$ the model implies perfect comovement between consumption, investment, and output, whereas $\omega = 0$ implies no decline in consumption in response to financial shocks; our choice of $\omega = 0.4$ implies that the model can closely replicate the fact that consumption drops by roughly one half of the amount of output in response to a financial shock.

4.4 Baseline Results

Figure (9) plots the impulse responses of the selected macroeconomic variables implied by the model in response to a financial shock, assuming the baseline specification of the monetary policy rule. The model captures remarkably well the shape of the corresponding impulse response functions shown in Figure (7). Consumption, investment, hours, and output all exhibit significant declines, with the peak response of each variable closely matching its empirical counterpart. While the model delivers hump-shaped dynamics for each of those variables that are consistent with those observed in the data, the model dynamics imply a

[24]In the case of no financial market frictions, $\chi = 0$. In this case, balance sheet conditions of the entrepreneurs are irrelevant for the cost of external funds and thus for their capital expenditure decisions.

peak response several quarters earlier than the peak response observed in Figure (7). The decline in the price level implied by the model also matches that seen in the data. Furthermore, given the estimated policy rule, the model-implied dynamics for inflation and output generate a path for the nominal interest rate that is broadly in line with the estimated response of the federal funds rate to an orthogonalized shock to the excess bond premium.

We now consider the ability of the model to explain the 2007–09 economic downturn as a response of the macroeconomy to a sequence of adverse financial shocks. To do so, we initialize the model to be in steady state as of the end of 2005 and feed in—as shocks to σ_t—the realized values of innovations to the excess bond premium over the 2006:Q1–2010:Q2 period. Figure (10) shows the evolution of the key macroeconomic variables of the U.S. economy over this period, while Figure (11) shows the corresponding path for the model-implied variables. According to the model, investment falls 14 percent in response to the realized shocks to the excess bond premium, a drop that is exactly in line with that observed in the data. The model-implied declines in hours and output are about 5 percent, which is slightly more than their empirical counterparts. Consistent with the data, the model also produces a modest—about 2 percent—but protracted decline in consumption.

Apart from its effects on the real economy, the model also implies that the realized sequence of shocks to the financial sector reduces inflation about 2 percentage points and causes a drop in the the nominal interest rate of more than 200 basis points. The response of inflation implied by the model is clearly smoother than that observed in the data, though their respective magnitudes are quite comparable. The actual response of the short-term nominal interest rate, however, is considerably stronger: The actual policy rate falls about 400 basis points, before leveling off at the zero lower bound.

4.5 Spread-Augmented Monetary Policy Rule

Overall, these results imply that disruptions to the financial intermediation process, as measured by the movements in the excess bond premium, can successfully account for the collapse in employment, consumption, investment and output during the recent financial crisis. Because the actual monetary policy response was clearly stronger than that implied by the model, we conjecture that the model is underestimating the full economic impact of financial shocks during the crisis. Nevertheless, the model dynamics are sufficiently close to the actual economic outcomes to provide a useful guide for alternative policy rules that may be used to stabilize the economy in the wake of disruptions in financial markets. We now consider one such rule proposed in the literature—namely, adjusting the first-difference rule so that monetary policy responds to financial distress as measured by the movements in credit spreads (cf. Cúrdia and Woodford [2009, 2010]).

In particular, we augment the baseline nominal interest rate rule by allowing for a direct response of the policy rate to the measured credit spread. Consistent with previous research, we set the coefficient on the spread equal to minus one, so that the nominal rate offsets the increase in financial stress on a one-for-one basis. The solid lines in Figure (12) depict the model impulse responses to a financial shock under the alternative monetary policy rule (i.e., $\phi_s = -1$), while the dotted lines denote the corresponding responses under the baseline rule (i.e., $\phi_s = 0$), replicated, for comparison purposes, from Figure (9).

The comparison of responses reveals that including the credit spread in the policy rule successfully stabilizes the real side of the economy. Importantly, the price level, in response to a financial shock, increases 0.2 percentage points, rather than falling 0.3 percentage points, as in the baseline case. Because of rising inflation, the nominal interest rate actually declines only 4 basis points under the spread-augmented rule, compared with the decline of 20 basis points in the baseline case. Effectively, the rise in inflation implies a reduction in the real rate of interest with very little movement in the nominal rate. This reduction in real rates leads to an offsetting increase in asset values and a much smaller decline in net worth than one sees in the baseline scenario. As a result, the response of the credit spread under the alternative policy basically mimics the response of the financial shock, resulting in very little additional amplification through the financial accelerator mechanism.

Although asset prices are forward looking, they influence the balance sheets of firms, and hence the strength of the financial accelerator, immediately. Consequently, a reduction in expected future real interest rates can be very effective in offsetting an emerging disruption in credit markets. (We return to this issue below, when considering policy responses in environments where financial disruptions are anticipated in advance.) This point is made explicit in Figure (13), which shows the model-implied path of the key macroeconomic aggregates during the 2006:Q1–2010:Q2 period under the spread-augmented policy rule. Consistent with the results presented in Figure (12), such a rule fully stabilizes the response of output and hours and implies only a modest reduction in investment—about 2 percentage points—which is offset by a slight increase in consumption. Again the response of the nominal interest rate is close to zero—the slight easing of monetary policy leads to an increase in inflation and a reduction in the real interest rate. These developments, in turn, boost asset values, counteracting the adverse consequences of financial shocks.

Although the disruptions in financial markets during the recent financial crisis are far more complex than simple shocks to the credit spread modelled in this framework, the above results nonetheless suggest that a monetary policy regime that is committed, in advance, to fully offset shocks to the financial system through active interest rate policy can be quite beneficial in mitigating the deleterious consequences of financial market disruptions. In the above examples, financial shocks are surprise events that imply an immediate jump in credit spreads. Because the output response is sluggish relative to the response of the credit spread, it is difficult to distinguish whether or not the benefits of responding to the movements in credit spreads are due to the fact that the widening of spreads signals a future decline in economic activity or that the agents in the economy anticipate the response and asset prices adjust accordingly.

To analyze this issue, we consider a shock to σ_t that is known four periods in advance. Thus, agents and the monetary authorities are able to anticipate future disruptions in credit markets. We again analyze the potential stabilization benefits of allowing the nominal interest rate to respond to the realized movements in credit spread. Figure (14) depicts the impulse responses to this anticipated financial shock under both the baseline (the dotted lines) and the spread-augmented policy rules (the solid lines). Under the baseline rule, the anticipated financial shock causes an immediate reduction in asset prices and entrepreneurial net worth and a decline in economic activity that occurs before the actual disruption in credit markets (i.e., period 4). The magnitude of these effects is only slightly less than that reported in response to an unanticipated shock in Figure (9).

In contrast, under the spread-augmented rule, the anticipated disruption to the credit intermediation process causes a slight initial decrease in the net worth of entrepreneurs, followed by a an additional modest decline upon the actual impact. The realized spread, however, reacts strongly once the shock occurs. In this case, the decline in output precede the spike in credit spreads. Hence, the stabilizing effects of the spread-augmented rule cannot be attributed to the fact that credit spreads provide a timely indicator of future cyclical downturns. Under this alternative policy, the decline in output is quite modest, relative to the case where policy does not respond to financial stress indicators. In addition, the announced policy of actively responding to credit spreads results in very little actual movement in the nominal interest rate. These results confirm the intuition suggested above that agents' expectations that the monetary authority will respond to financial disruptions when they occur can have a powerful stabilizing effects on economic activity.

5 Conclusion

This paper examines the extent to which the workhorse DSGE model with financial frictions is able to account for the extraordinary macroeconomic dynamics of the 2007–09 financial period. We show that by carefully constructing a sequence of financial shocks, a reasonably calibrated version of the CEE/SW framework augmented with the BGG financial accelerator can fully account for the overall drop in consumption, investment, hours, and output that was observed during the crisis period. The model also does well at matching the observed decline in inflation and nominal interest rates, as well as the sharp widening of nonfinancial credit spreads. Although the model is relatively simple compared with the recent work in this area, these results nonetheless provide considerable insight into the importance of financial factors in business cycle fluctuations. In particular, our results suggest that by allowing the nominal interest rate to respond to credit spreads, monetary policy can effectively dampen the negative consequences of financial market shocks on real economic activity, while experiencing only a modest increase in inflation.

Tables and Figures

Table 1: Summary Statistics of Financial Corporate Bonds

Variable	Mean	SD	Min	P50	Max
No. of bonds per firm/month	3.00	3.46	1.00	2.00	26.0
Mkt. value of issue[a] ($mil.)	466.9	552.6	9.1	264.0	4,350
Maturity at issue (years)	10.4	8.0	2.0	10.0	40.0
Term to maturity (years)	8.6	7.7	1.0	5.9	30.0
Duration (years)	6.47	3.15	0.90	4.79	15.3
Credit rating (S&P)	-	-	CC	A2	AAA
Coupon rate (pct.)	6.89	1.94	2.25	6.63	15.75
Nominal effective yield (pct.)	6.78	2.77	1.01	6.43	41.2
Credit spread (bps.)	172	253	5	106	3,495

NOTE: Sample period: Jan1985–June2010; Obs. = 42,880; No. of bonds = 886; No. of firms = 193. 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).

Table 2: Estimates of the Bond-Pricing Model by Type of Bond

	Non-callable		Callable	
Explanatory Variable	*Est.*	*S.E.*	*Est.*	*S.E.*
$-DD_{it}$	0.180	0.035	0.159	0.019
$(-DD_{it})^2$	0.007	0.002	0.005	0.001
$\ln(DUR_{it}[k])$	0.325	0.029	-0.027	0.047
$\ln(PAR_{it}[k])$	0.171	0.026	0.009	0.034
$\ln(CPN_i[k])$	0.295	0.087	1.073	0.154
LEV_t	-	-	-0.455	0.065
SLP_t	-	-	-0.175	0.028
CRV_t	-	-	-0.047	0.031
VOL_t	-	-	0.076	0.009
Adj. $R^2 = 0.626$				
$\Pr > W_R$:[a] 0.000				
$\Pr > W_I$:[b] 0.000				

NOTE: Sample period: Jan1985–Jun2010; Obs. = 42,880; No. of bonds/firms = 886/193. Dependent variable in the bond-pricing regression is $\ln(S_{it}[k])$, the logarithm of the credit spread on bond k (issued by firm i) in month t. The regression includes a constant term (not reported) and is estimated by OLS (see text for details). Robust asymptotic standard errors are double clustered in the firm (i) and time (t) dimensions; see Cameron et al. [2010] for details.

[a] p-value for the Wald test of the exclusion of credit rating fixed effects.

[b] p-value for the Wald test of the exclusion of industry fixed effects.

Table 3: Financial Indicators and the Excess Bond Premium

Financial Indicator	(1)[a]	(2)[b]	(3)	(4)	(5)[c]	(6)[d]
S&P 100 Volatility (VXO)	0.005	-	-	-	-	-
	(0.004)					
Eurodollar volatility (3m)	-	0.030	-	-	-	-
		(0.017)				
Treas. volatility (30y)	-	-	0.015	-	-	-
			(0.011)			
Libor–Treas. (3m)	-	-	-	0.078	-	-
				(0.058)		
Swap–Treas. (5y)	-	-	-	-	0.558	-
					(0.220)	
Off/On-the-run Treas. (10y)	-	-	-	-	-	-0.000
						(0.005)
Adj. R^2	0.620	0.630	0.622	0.620	0.652	0.627
Obs.	293	257	305	305	259	268

NOTE: Sample period: Jan1985–Jun2010, unless noted otherwise. Dependent variable in the regression is EBP_{t+1}, the excess financial bond premium in month $t+1$. In addition to the specified financial indicators in month t, each specification also includes a constant and EBP_t (not reported). Entries in the table denote OLS estimates of the coefficients associated with each financial indicator; standard errors reported in parentheses are based on the asymptotic covariance matrix computed according to Newey and West [1987], with the "lag truncation" parameter $L = 12$.

[a]Sample period: Jan1986–Jun2010.

[b]Sample period: Jan1989–Jun2010.

[c]Sample period: Nov1988–Jun2010.

[d]Sample period: Feb1988–Jun2010.

Table 4: Excess Bond Premium and Real GDP

Financial Indicator	Forecast Horizon: 1 quarter	Forecast Horizon: 4 quarters
Term spread (3m–10y)	-0.193	-0.351
	[1.75]	[3.14]
Real FFR	0.253	0.220
	[1.88]	[1.70]
Excess bond premium	-0.500	-0.528
	[3.34]	[5.08]
Adj. R^2	0.392	0.336
Obs.	101	98

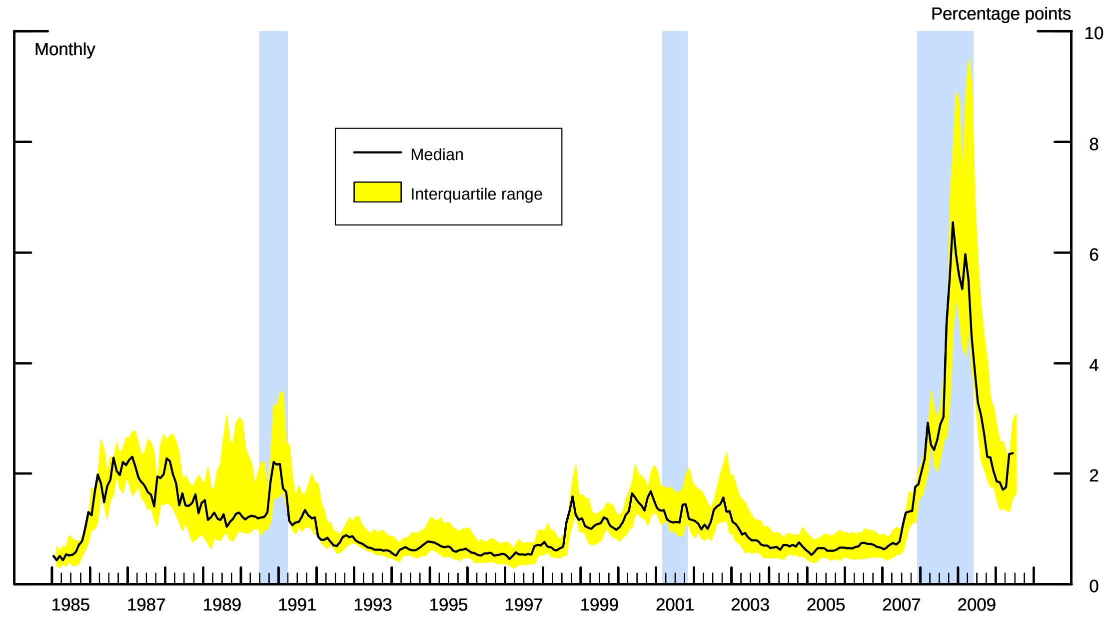
NOTE: Sample period: 1985:Q1–2010:Q2. Dependent variable is $\Delta^h Y_{t+h}$, where Y_t denotes the log of real GDP in quarter t and h is the forecast horizon. In addition to the specified financial indicator in quarter t, each specification also includes a constant, current, and p lags of ΔY_t (not reported), where p is determined by the AIC. Entries in the table are the standardized estimates of the OLS coefficients associated with each financial indicator. For the 1-quarter horizon, absolute t-statistics reported in brackets are based on the asymptotic covariance matrix (HC3) computed according MacKinnon and White [1985]; for the 4-quarter horizon, absolute t-statistics are computed according to Hodrick [1992].

Table 5: Excess Bond Premium and Components of Aggregate Demand

	Forecast Horizon: 1 quarter						
Financial Indicator	C-NDS	C-D	I-RES	I-ES	I-HT	I-NRS	INV
Term spread	-0.163	-0.070	-0.169	-0.042	-0.092	0.282	-0.112
	[1.61]	[0.53]	[1.96]	[0.47]	[0.75]	[2.38]	[1.28]
Real FFR	0.095	0.027	0.055	0.046	0.057	-0.009	0.142
	[0.68]	[0.19]	[0.59]	[0.37]	[0.43]	[0.06]	[1.40]
Excess bond premium	-0.239	-0.438	0.039	-0.642	-0.217	-0.265	-0.372
	[2.20]	[3.45]	[0.28]	[4.31]	[2.20]	[2.08]	[5.63]
Adj. R^2	0.429	0.137	0.509	0.422	0.300	0.321	0.516
Obs.	101	101	101	101	101	101	101
	Forecast Horizon: 4 quarters						
Financial Indicator	C-NDS	C-D	I-RES	I-ES	I-HT	I-NRS	INV
Term spread (3m–10y)	-0.307	-0.247	-0.507	-0.228	-0.193	0.509	-0.193
	[3.18]	[1.24]	[5.82]	[2.32]	[1.59]	[3.41]	[2.58]
Real FFR	0.185	0.093	0.268	-0.050	0.017	-0.174	0.127
	[1.40]	[0.39]	[2.75]	[0.37]	[0.13]	[1.04]	[1.28]
Excess bond premium	-0.316	-0.330	0.095	-0.738	-0.236	-0.677	-0.668
	[3.29]	[1.86]	[1.06]	[4.47]	[2.19]	[4.73]	[7.56]
Adj. R^2	0.470	0.099	0.486	0.453	0.352	0.656	0.601
Obs.	98	98	98	98	98	98	98

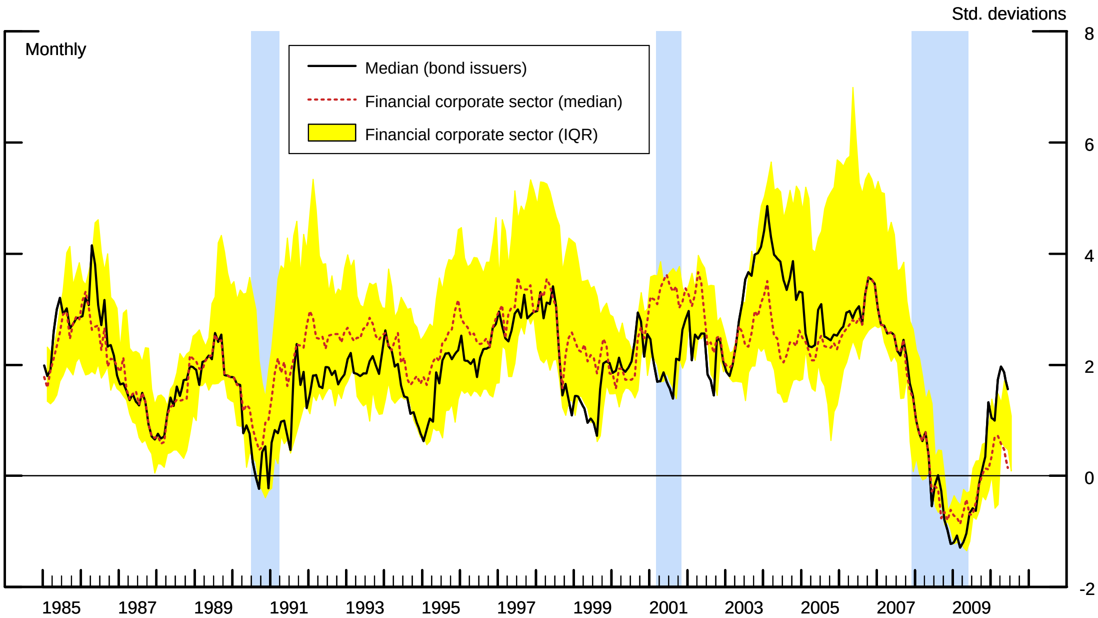
NOTE: Sample period: 1985:Q1–2010:Q2. Dependent variable is $\Delta^h Y_{t+h}$, where Y_t denotes the log of the component of private (real) aggregate demand in quarter t and h is the forecast horizon: C-D = PCE on durable goods; C-NDS = PCE on nondurable goods & services; I-RES = residential investment; 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, current, and p lags of ΔY_t (not reported), where p is determined by the AIC. Entries in the table are the standardized estimates of the OLS coefficients associated with each financial indicator. For the 1-quarter horizon, absolute t-statistics reported in brackets are based on the asymptotic covariance matrix (HC3) computed according MacKinnon and White [1985]; for the 4-quarter horizon, absolute t-statistics are computed according to Hodrick [1992].

Figure 1: Financial Corporate Credit Spreads



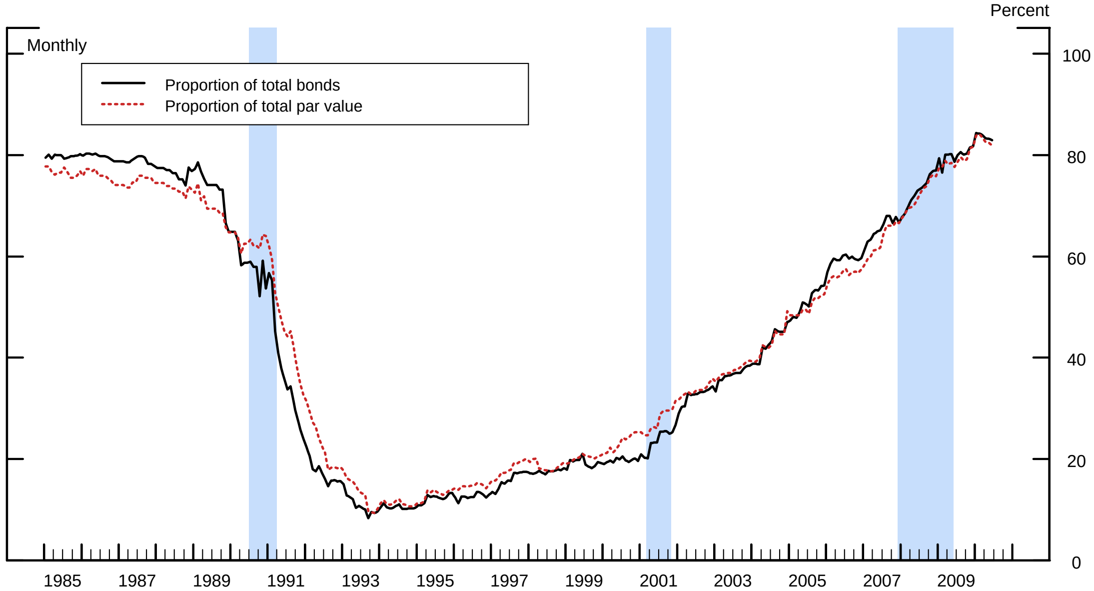
NOTE: Sample period: Jan1985–Jun2010. The solid line depicts the median spread on senior unsecured bonds issued by financial firms in our sample, and the shaded band depicts the corresponding interquartile range. The shaded vertical bars denote the NBER-dated recessions.

Figure 2: Distance-to-Default in the Financial Corporate Sector



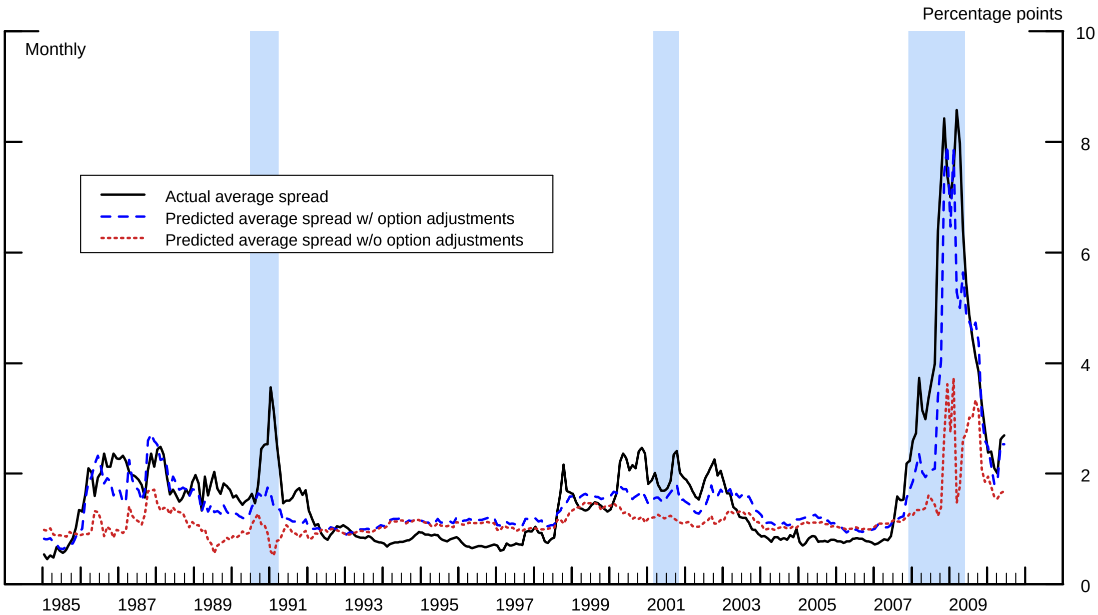
Note: Sample period: Jan1985–Jun2010. The solid line depicts the (weighted) median distance-to-default (DD) for our sample of bond issuers. The dotted line depicts the (weighted) median DD for the U.S. financial corporate sector, and the shaded band depicts the corresponding interquartile range. The firm-specific year-ahead DDs are calculated using the Merton [1974] model (see text for details); all percentiles are weighted by firm liabilities. The shaded vertical bars denote the NBER-dated recessions.

Figure 3: Callable Financial Corporate Bonds



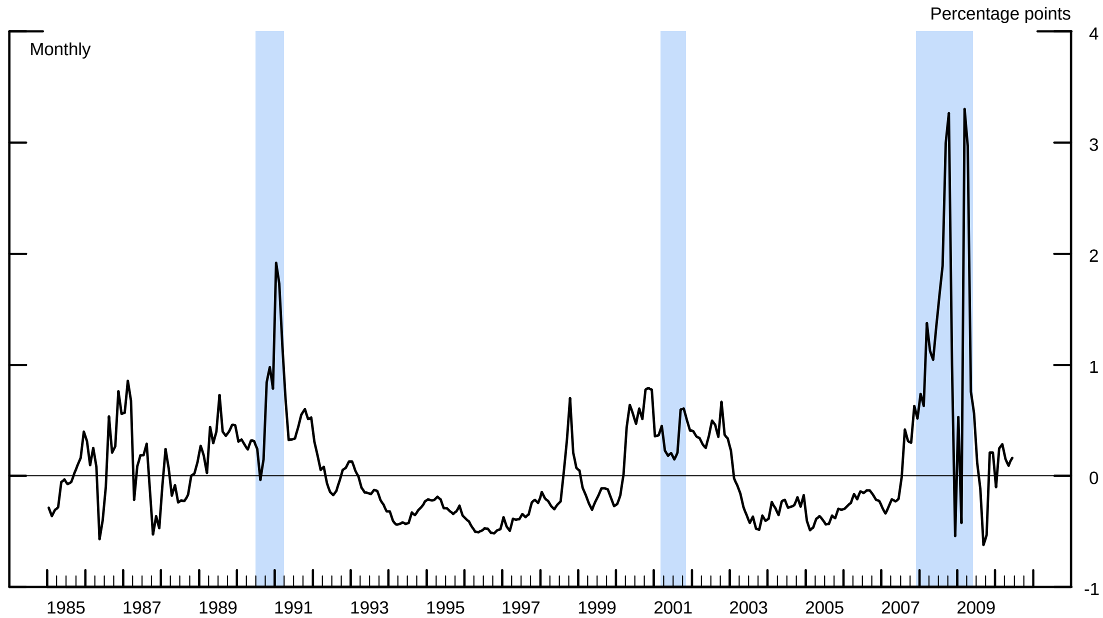
NOTE: Sample period: Jan1985–Jun2010. The figure depicts the proportion of bonds in our sample that are callable. The shaded vertical bars denote the NBER-dated recessions.

Figure 4: Actual and Predicted Financial Corporate Credit Spreads



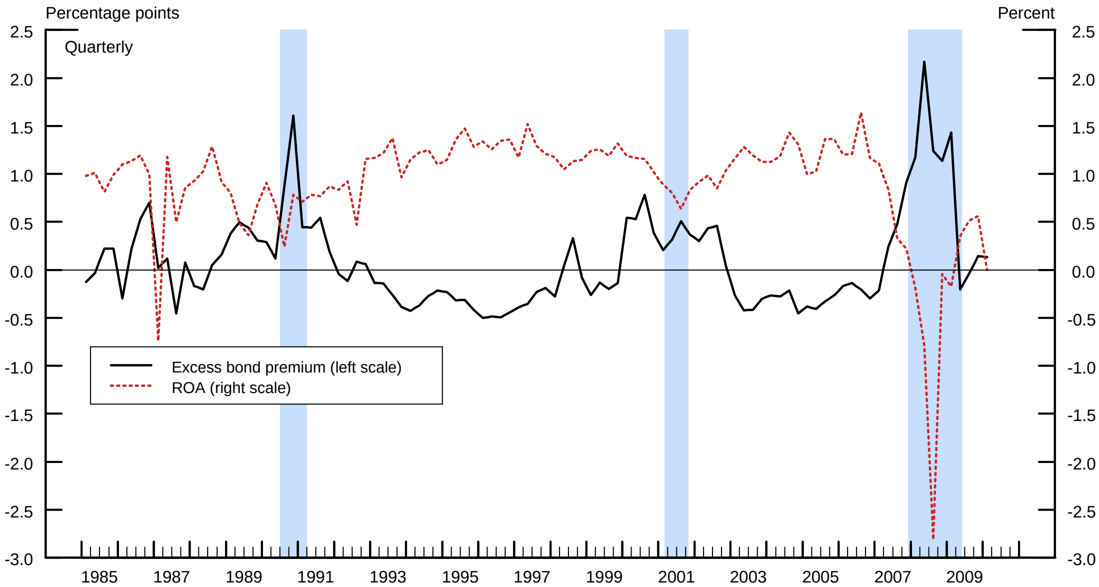
NOTE: Sample period: Jan1985–Jun2010. The solid line depicts the average credit spread on senior unsecured bonds issued by financial firms in our sample. The dashed line depicts the predicted average credit spread based on the bond-pricing model that includes the option-adjustment terms; the dotted line depicts the predicted average credit spread based on the bond-pricing model that excludes the option-adjustment terms (see text for details). The shaded vertical bars denote the NBER-dated recessions.

Figure 5: Excess Bond Premium in the Financial Corporate Sector



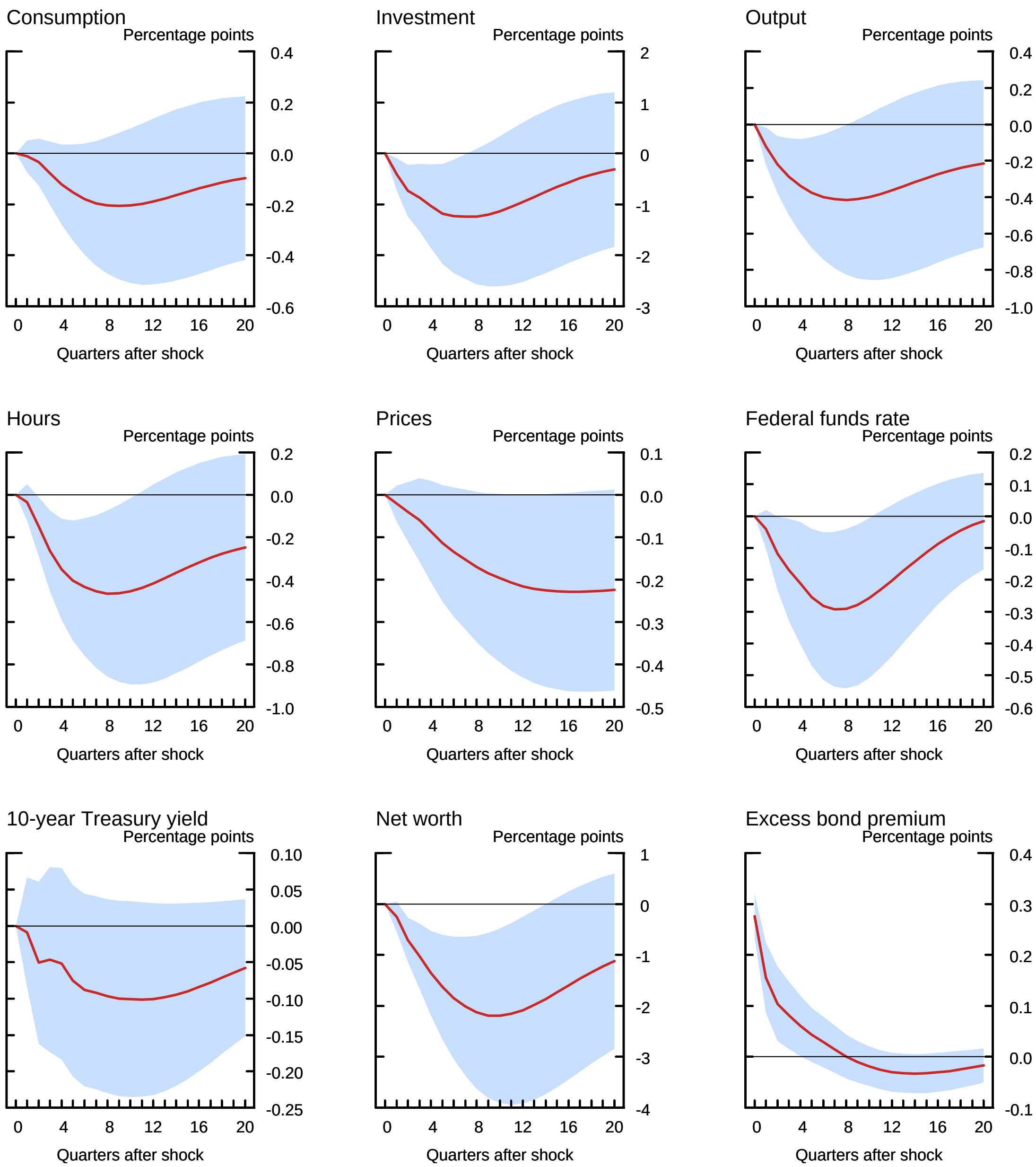
NOTE: Sample period: Jan1985–Jun2010. The figure depicts the estimated (option-adjusted) excess bond premium in the U.S. financial sector (see text for details). The shaded vertical bars denote the NBER-dated recessions.

Figure 6: Excess Bond Premium and Financial Sector Profitability



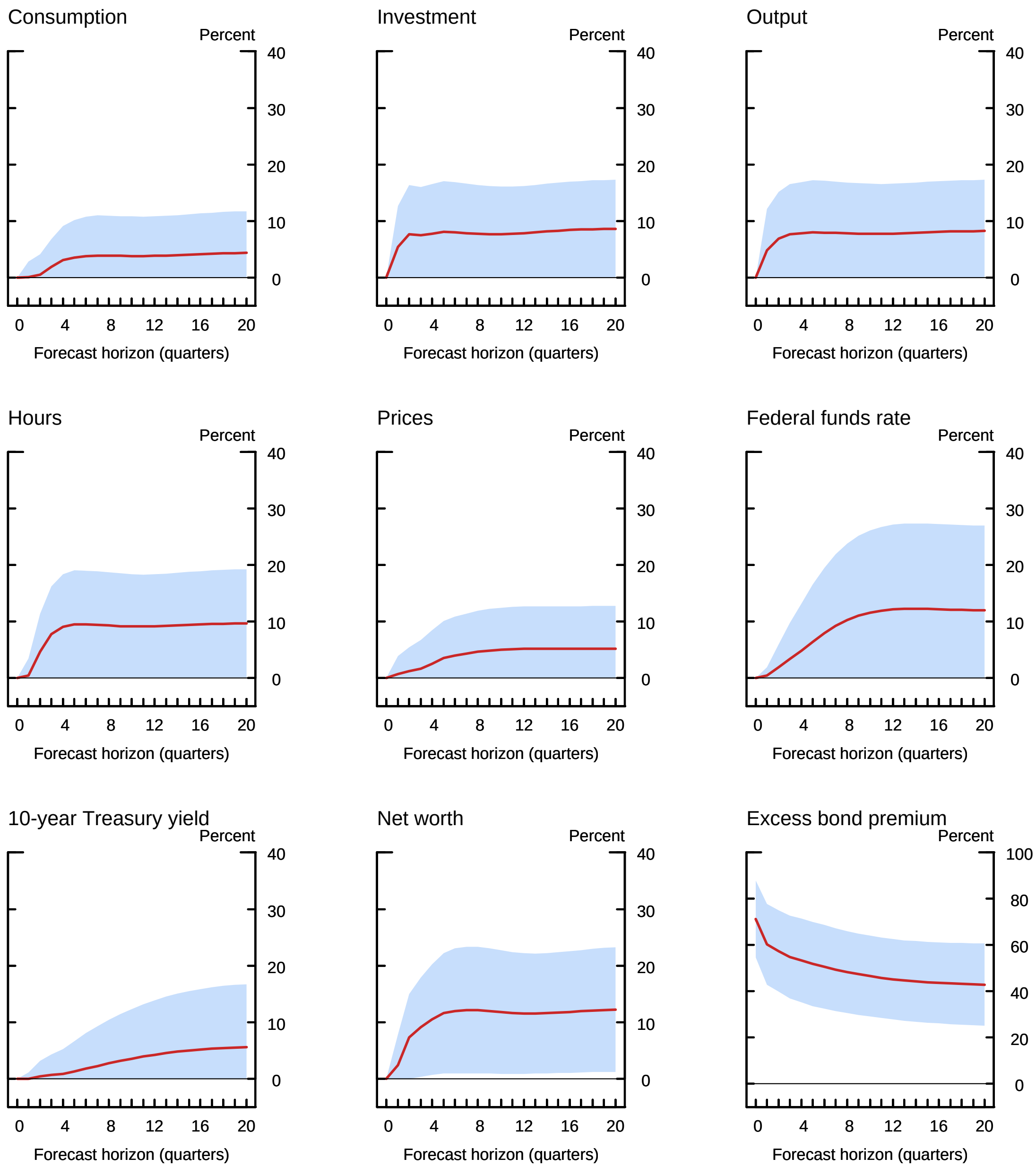
NOTE: Sample period: 1985:Q1–2010:Q2. The solid line depicts the quarterly average of the estimated (option-adjusted) excess bond premium (see text for details). The dotted line depicts the annualized return on assets (ROA) for the U.S. financial corporate sector. The shaded vertical bars denote the NBER-dated recessions.

Figure 7: Macroeconomic Implications of a Financial Shock



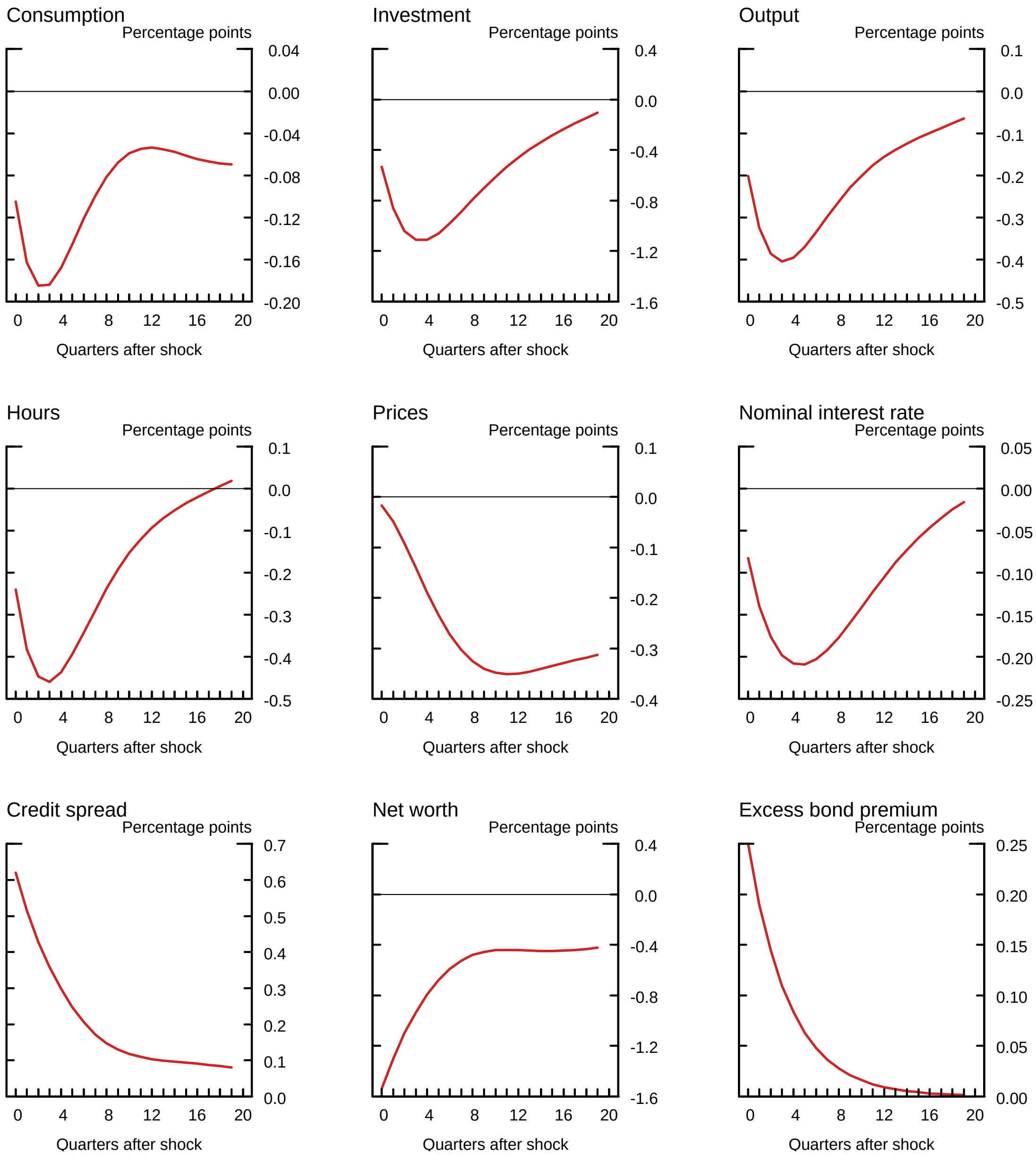
NOTE: The figure depicts the impulse response functions from a 9-variable VAR(2) model to a 1 standard deviation orthogonalized shock to the excess financial bond premium (see text for details). Shaded bands denote 95-percent confidence intervals based on 1,000 bootstrap replications.

Figure 8: Forecast Error Variance Decomposition



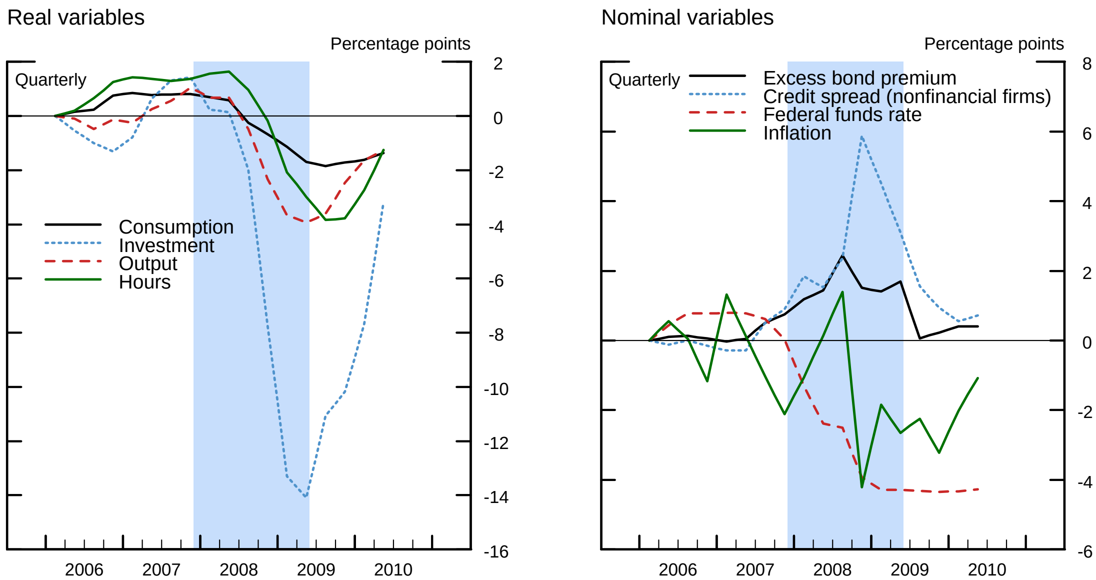
NOTE: The figure depicts the forecast error variance decomposition from a 9-variable VAR(2) model to a 1 standard deviation orthogonalized shock to the excess financial bond premium (see text for details). Shaded bands denote 95-percent confidence intervals based on 1,000 bootstrap replications.

Figure 9: Model-Based Impulse Responses to a Financial Shock
(Baseline Monetary Policy Rule)



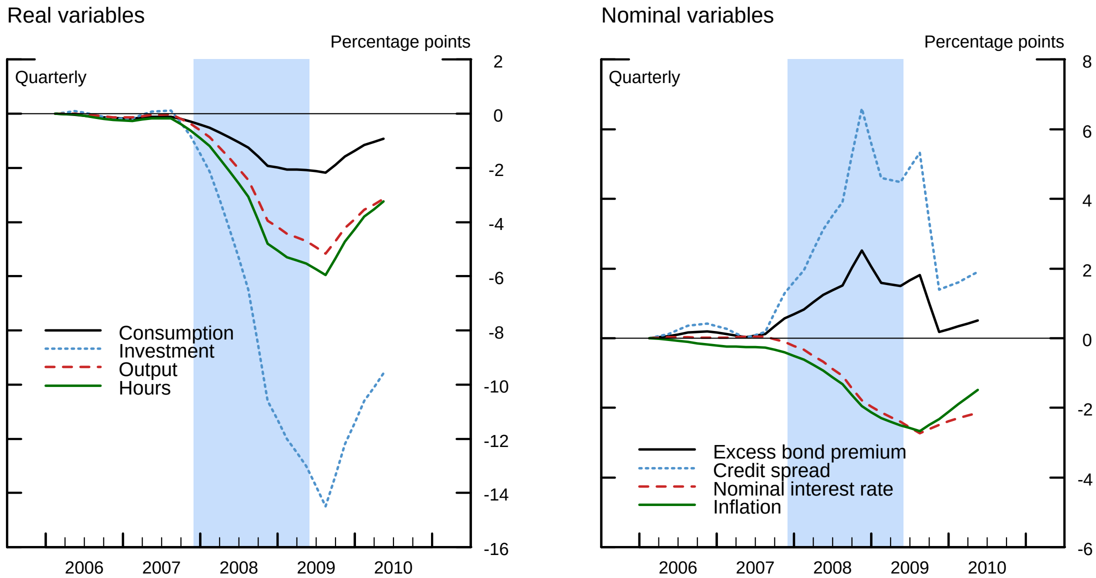
Note: The panels of the figure depict the model-based impulse response functions of selected variables to a 1 standard deviation financial shock for the baseline specification of the monetary policy rule, a case in which the monetary authority does not respond to credit spreads (see text for details). All variables are expressed in percentage-point deviations from their respective steady-state values.

Figure 10: The U.S. Economy During the 2007–09 Financial Crisis



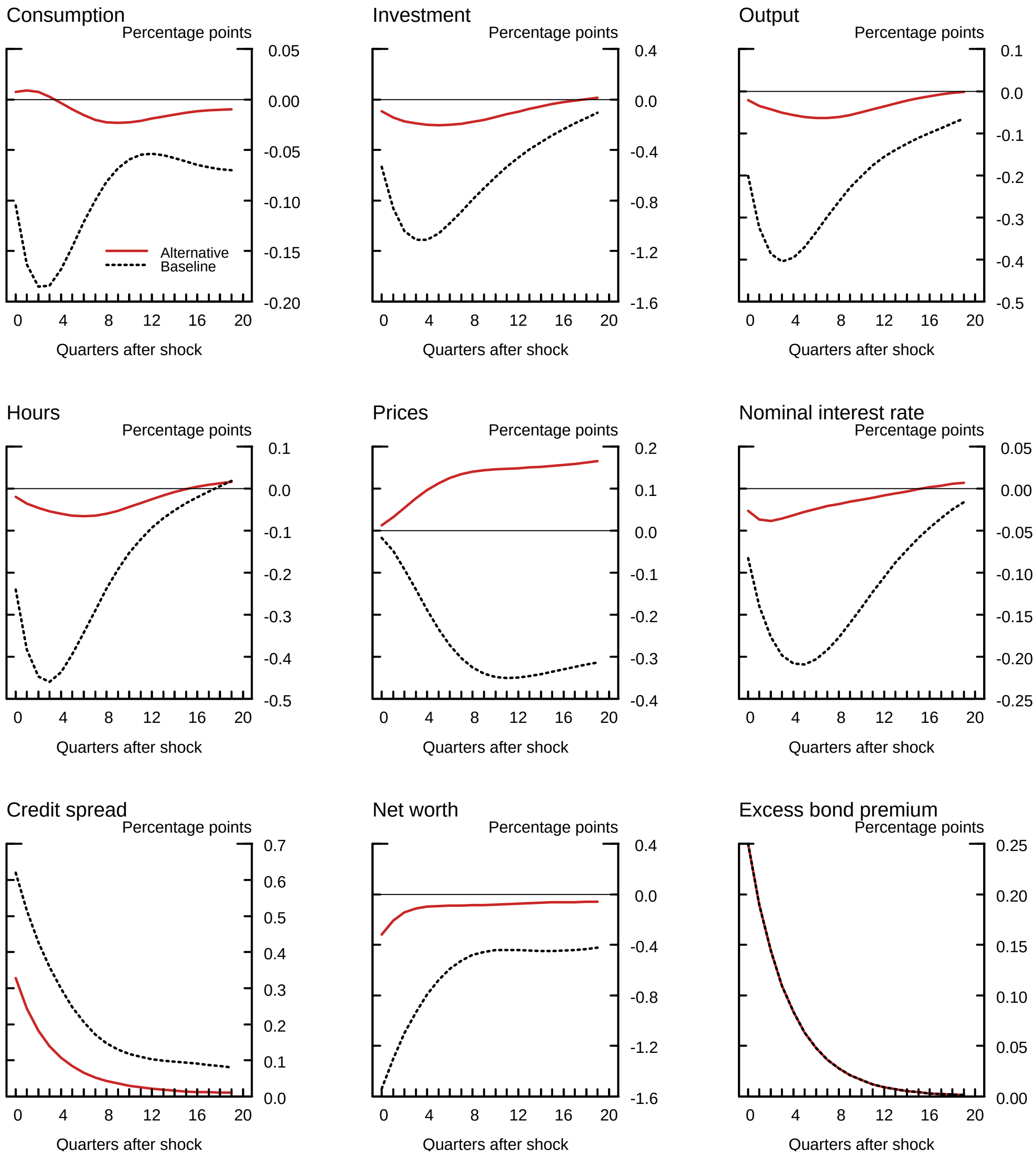
NOTE: The two panels of the figure depict the behavior of selected macroeconomic variables in the period surrounding the 2007–09 financial crisis. Consumption, investment, output, and hours worked have been detrended using the Hodrick-Prescott filter. All series are indexed to equal zero in 2006:Q1. The shaded vertical bar denotes the NBER-dated recession.

Figure 11: Model-Based Simulation of a Financial Shock
(Baseline Monetary Policy Rule)



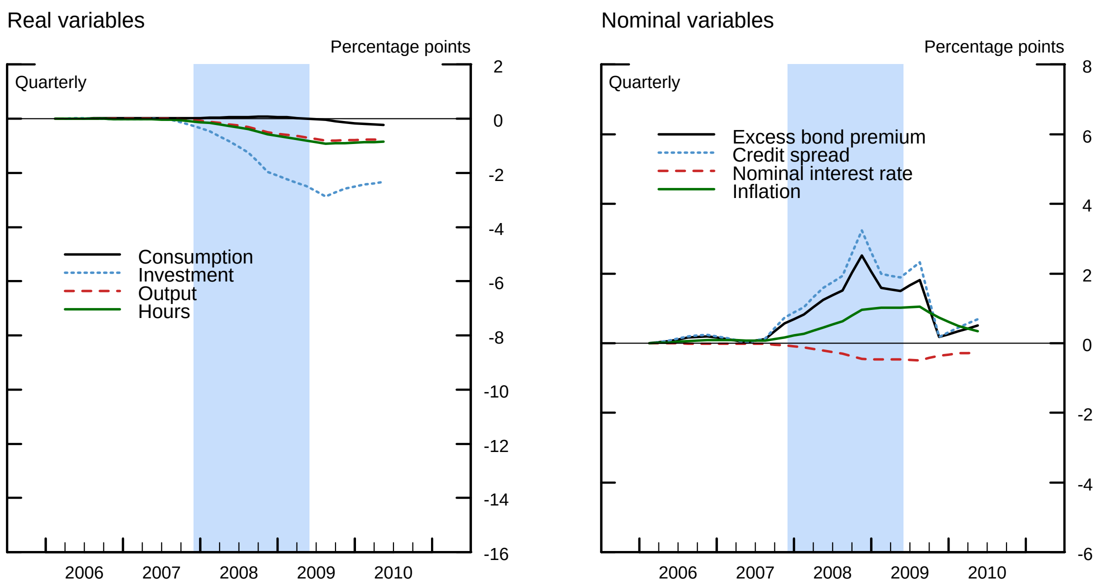
NOTE: The two panels of the figure depict the model-implied path of selected macroeconomic variables in response to the estimated financial shocks for the baseline specification of the monetary policy rule, a case in which the monetary authority does not respond to credit spreads (see text for details). All variables are expressed in percentage-point deviations from their respective steady-state values. The shaded vertical bar denotes the NBER-dated recession.

Figure 12: Model-Based Impulse Responses to a Financial Shock
(Baseline vs. Spread-Augmented Monetary Policy Rule)



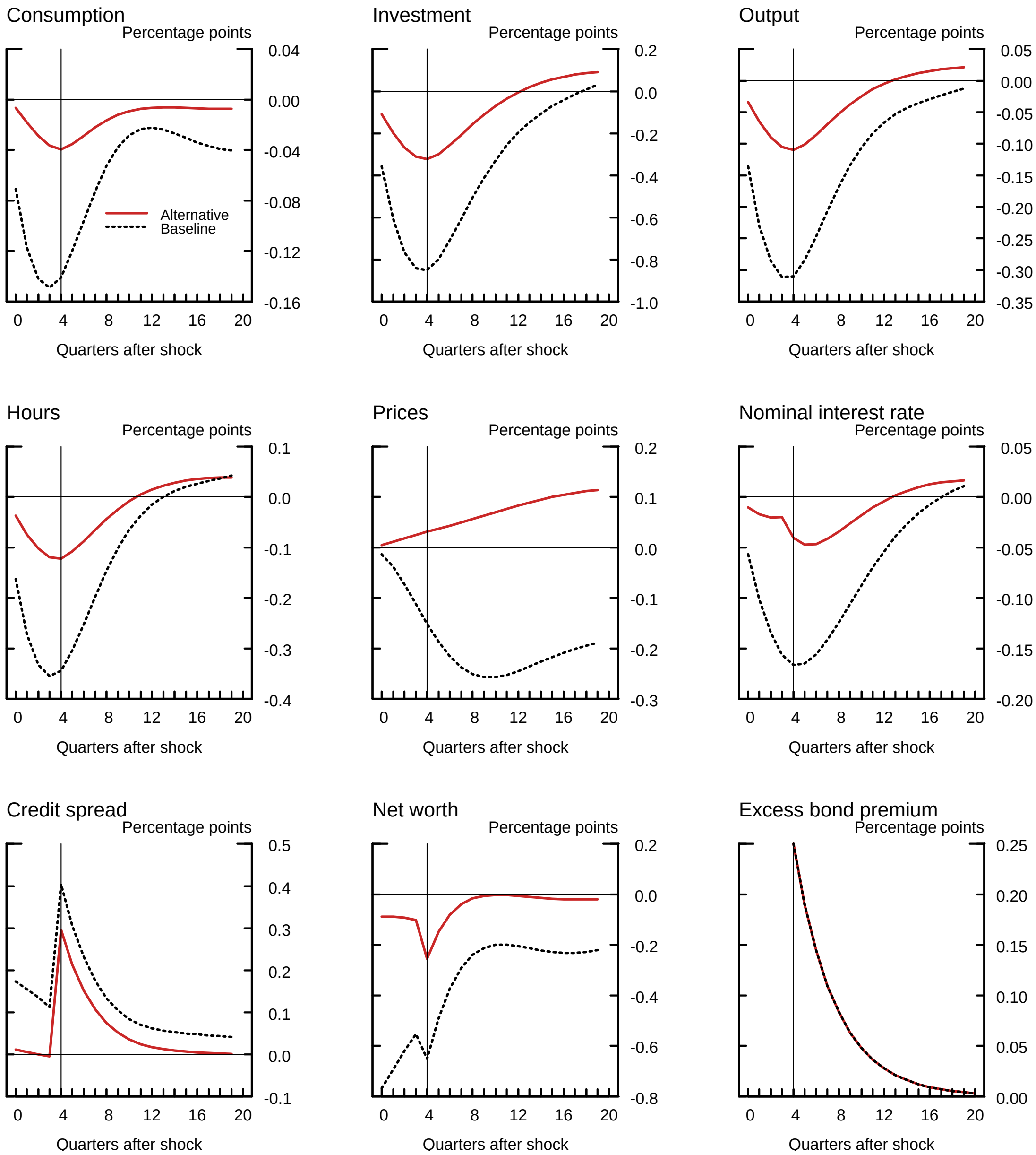
NOTE: The solid lines in each panel of the figure depict the model-based impulse response functions of selected variables to a 1 standard deviation financial shock for the alternative specification of the monetary policy rule, a case in which the monetary authority responds to credit spreads; the dotted lines in each panel correspond to impulse responses under the baseline specification of the monetary policy rule (see text for details). All variables are expressed in percentage-point deviations from their respective steady-state values.

Figure 13: Model-Based Simulation of a Financial Shock
(Spread-Augmented Monetary Policy Rule)



NOTE: The two panels of the figure depict the model-implied path of selected macroeconomic variables in response to the estimated financial shocks for the alternative specification of the monetary policy rule, a case in which the monetary authority responds to credit spreads (see text for details). All variables are expressed in percentage-point deviations from their respective steady-state values. The shaded vertical bar denotes the NBER-dated recession.

Figure 14: Model-Based Impulse Responses to an Anticipated Future Financial Shock (Baseline vs. Spread-Augmented Monetary Policy Rule)



Note: The solid lines in each panel of the figure depict the model-based impulse response functions of selected variables to an anticipated 4-quarter-ahead 1 standard deviation financial shock for the alternative specification of the monetary policy rule, a case in which the monetary authority responds to credit spreads; the dotted lines in each panel correspond to impulse responses under the baseline specification of the monetary policy rule (see text for details). All variables are expressed in percentage-point deviations from their respective steady-state values.

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