

The Decline of Too Big to Fail

PRELIMINARY DRAFT

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May 2019

Abstract

For globally systemically important banks (G-SIBs) with U.S. headquarters, we find large post-Lehman reductions in market-implied probabilities of government bailout, along with big increases in debt financing costs for these banks after controlling for insolvency risk. The data are consistent with significant effectiveness for the official sector’s post-Lehman G-SIB failure-resolution intentions, laws, and rules. G-SIB creditors now appear to expect to suffer much larger losses in the event that a G-SIB approaches insolvency. In this sense, we estimate a major decline of “too big to fail.”

JEL Classifications: G12, G13, G22, G24

Keywords: Too big to fail, systemically important banks, government bailouts

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1. Introduction

Crisis revelations of the costs of “too-big-to-fail” have led to new legal methods, globally, for resolving the insolvencies of systemically important banks. Rather than bailing out these firms with government capital injections, insolvency losses are now supposed to be allocated to wholesale creditors. As a consequence, major credit rating agencies have substantially reduced or removed explicit “sovereign uplifts” to the ratings of the senior unsecured debt of the holding companies of U.S. globally systemically important banks (G-SIBs).

Many market participants believe, however, that these reforms have not eliminated the likelihood of government bailouts of these firms.¹ Our main objective is to estimate post-crisis declines in market-implied bailout probabilities, the associated increases in G-SIB bond yields, and the declines in G-SIB equity market values stemming from reductions in debt financing subsidies associated with bailout expectations.

We show that G-SIB balance sheet data and the market prices of debt and equity imply a dramatic and persistent post-crisis reduction in market-implied probabilities of government bailouts of U.S. G-SIB holding companies. We also report similar but smaller effects for domestically important non-G-SIB banks, or “D-SIBs.”² Our sample period is 2002-2017. Our demarcation point for measuring a change in bailout probabilities is the bankruptcy of Lehman Brothers in September 2008. We refer to the prior period as “pre-Lehman” and the subsequent period as “post-Lehman.” Many market participants were surprised that the U.S. government did not bail out Lehman.³ We cannot disentangle how much of the post-Lehman reduction in investor bailout expectations is due to the effectiveness of new failure resolution methods, which have not yet been tried in practice, as opposed to a post-Lehman updating of beliefs about government bailout preferences.

Our results are based in part on a large-scale panel analysis of corporate credit spreads observed in the market for the credit default swaps (CDS) of over 630 public U.S. firms. Of our sample of firms, only large banks are assumed to have had a significant change in bailout probabilities. Our central

¹See [Government Accountability Office \(2014\)](#).

²European regulators officially designate their domestically systemically important banks (D-SIBs). While there is no such official designation by U.S. regulators, we label as a “D-SIB” any publicly traded bank that is not a G-SIB and is required to undergo stress tests under the Comprehensive Capital Analysis and Review (CCAR) and the Dodd-Frank Act stress test (DFAST).

³See, for example, the New York Times article “Revisiting the Lehman Brothers Bailout That Never Was” by [Stewart and Eavis \(2018\)](#).

measures of the credit worthiness of a firm are its distance to default and the expected fractional loss of bond notional given no bailout at insolvency. Conceptually, the distance to default of a firm is the number of standard deviations of annual changes in its asset value by which the current asset value exceeds the insolvency level of assets. Distance to default is thus a risk-adjusted measure of a firm’s capital buffer, and a strong predictor of default.⁴ We refer to the loss of bond notional given no bailout at insolvency as the loss given liquidation (LGL). For a given liability structure, variation in LGL is closely linked with variation in the (risk-adjusted) distance between book liabilities and the default threshold.

Under idealized theoretical conditions, for a given risk-neutral probability of no bailout, credit spreads are explained by distance to default, default risk premia and LGL. For a given risk-neutral insolvency probability and LGL, credit spreads are essentially proportional to the risk-neutral probability of no bailout. For example, if the no-bailout probability of firm *A* is twice that of firm *B* then the credit spread of firm *A* is twice that of firm *B*, all else equal. We detect large post-Lehman changes in the risk-neutral (or “market-implied”) probability of bank bailout based on changes in the observed relationship between credit spreads, distance to default and loss given liquidation. This change is far larger than that associated with the general post-Lehman increase in market premia for bearing corporate default risk. By including a large number of non-banks in our sample, we control for variation over time in corporate default risk premia, which is substantial (Berndt, Douglas, Duffie, and Ferguson, 2018). Rather than taking default risk premia for non-banks and banks to be the same, we rely only on the assumption that the ratio of their default risk premia did not change with the post-Lehman change in bailout probabilities.

Bailout expectations are a key input to the measurement of a firm’s distance to default. For example, if at the end of December 2017 one changes one’s assumption of Goldman Sachs’ bailout probability from 0.6 to 0.2 while holding fixed its observed balance sheet, the fitted distance to default of Goldman Sachs increases by about two standard deviations. This effect is measured using a structural dynamic model of debt and equity prices that extends Leland (1994a). At a modeled bailout, the government injects enough capital to return the balance sheet of the bank to a given “safe” condition. At any time before insolvency, the equilibrium prices of debt and equity reflect consistent expectations regarding the

⁴See Duffie (2011) for a summary of relevant research.

likelihood of bailout and the post-bailout capital structure of the bank. These post-bailout continuation values include the market values of subsidies associated with successive future potential bailouts. For instance, on average in the pre-Lehman period, we find that one-half to two-thirds of the market value of future bailout subsidies is associated with the next bailout, with the remainder of the subsidy value being associated with subsequent bailouts.

We face the following identification challenge. The impact on debt and equity prices of a downward shift in default risk premia can be approximately offset by some common upward shift in assumed pre-Lehman and post-Lehman no-bailout probabilities. So, our approach is not to provide point estimates of *both* a pre-Lehman bailout probability π_{pre} *and* a post-Lehman bailout probability π_{post} , but rather to estimate a schedule of pairs $(\pi_{\text{pre}}, \pi_{\text{post}})$ that are jointly consistent with the data. For example, when G-SIBs are assumed to have a post-Lehman bailout probability of $\pi_{\text{post}} = 0.2$, we estimate a pre-Lehman bailout probability π_{pre} of 0.63. Alternatively, for a post-Lehman bailout probability π_{post} of 0.3, our estimate of π_{pre} is 0.67. Schedules of estimated pairs of pre-Lehman and post-Lehman bailout probabilities are tabulated later in the paper. When allowing for heterogeneity across banks, we find substantial cross-sectional variation in bailout probabilities.

We find that D-SIBs have smaller post-Lehman declines in bailout probabilities than G-SIBs. This is natural, given that D-SIBs are by definition not as big as G-SIBs and thus less likely to be viewed by regulators and creditors as too big to fail. For example, for a post-Lehman D-SIB bailout probability of 0.2, the data imply a pre-Lehman D-SIB bailout probability of 0.46, which is much lower than the associated G-SIB pre-Lehman estimated bailout probability of 0.63.

Given our estimated post-Lehman reductions in bailout probabilities, we quantify the associated reductions in effective government subsidies provided to the shareholders of G-SIBs through lower debt financing costs. For example, we estimate that about 29% of the equity market value of G-SIBs, on average during the pre-Lehman period, can be ascribed to bailout-subsidized debt financing costs. This is based on a hypothetical reduction in the G-SIB bailout probability from the estimated pre-Lehman level of 0.63 to the assumed post-Lehman level of 0.2. At a fixed distance to default and loss given liquidation, this reduction in bailout probability implies post-Lehman senior unsecured yield spreads that are roughly twice what they would have been had there been no post-Lehman decline in bailout probability.

The remainder of the paper is structured as follows. Section 2 reviews closely related prior research. Section 3 is a high-level preview of our empirical identification strategy. Section 4 reviews a basic theoretical model of the valuation of a bank’s debt and equity that allows for bailout. Among other purposes, this model captures the effect of a given assumed bailout probability on the measured distance to default and loss given liquidation of a bank, which are the key inputs to our empirical estimation of bailout probabilities. Section 5 describes the data and presents descriptive statistics. Section 6 explains how we choose the model parameters that capture the influence of bailout on a big bank’s measured distance to default and LGL. Section 7 presents our estimates of bailout probabilities and their implications for big-bank credit spreads and equity subsidies, among other effects. Section 8 discusses alternatives to our main hypothesis and then concludes. The appendices include additional details, both theoretical and empirical.

2. Prior Related Work

Of the large empirical literature on too-big-to-fail (TBTF) subsidies,⁵ relatively few studies address the degree to which there has been a post-crisis decline in TBTF subsidies. None of these studies estimate post-crisis changes in bailout probabilities.

Of the body of prior research on the post-crisis decline of TBTF, the closest point of comparison

⁵Years before the Great Recession, [Stern and Feldman \(2004\)](#) stressed the importance of the TBTF problem, arguing that a safety net provided by the government lowers creditors incentives to monitor and banks’ incentive to act prudently. [Mishkin \(2006\)](#), however, argued that [Stern and Feldman \(2004\)](#) overstated the importance of the TBTF problem. Using international data, [Mäkinen, Sarno, and Zinna \(2018\)](#) find a risk premium associated with implicit government guarantees. They suggest that the risk premium is tied to sovereign risk, meaning guaranteed banks inherit guarantors risk. [Gandhi, Lustig, and Plazzi \(2016\)](#) also provide empirical evidence consistent with the idea that stock-market investors price in the implicit government guarantees that protect shareholders of the largest banks in developed countries. [Minton, Stulz, and Taboada \(2017\)](#), on the other hand, find no evidence that large banks are valued more highly than other firms. [O’Hara and Shaw \(1990\)](#) find positive wealth effects accruing to TBTF banks, with corresponding negative effects accruing to less systemically important banks. [Kelly, Lustig, and Nieuwerburgh \(2016\)](#) use options data to show that a collective government guarantee for the financial sector lowers index put prices far more than those of individual banks and explains the increase in the basket-index put spread observed during the Great Financial Crisis. [Schweikhard and Tsesmelidakis \(2012\)](#) investigate the impact of government guarantees on the pricing of default risk in credit and stock markets and, using a Merton-type credit model with exogenous default boundary, provide evidence of a structural break in the valuation of U.S. bank debt in the course of the 2007–2009 financial crisis, manifesting in a lowered default boundary, or, under the pre-crisis regime, in higher stock-implied credit spreads. [Balasubramnian and Cyree \(2011\)](#) claim that the TBTF discount on yield spreads is absent prior to the Long-Term Capital Management (LTCM) bailout. They find a paradigm shift in determinants of yield spreads after the LTCM bailout. [Santos \(2014\)](#) demonstrates that the additional discount that bond investors offer the largest banks compared with the return they demand from the largest non-banks and non-financial corporations is consistent with the idea that investors perceive the largest U.S. banks to be too big to fail. The impact of subsidies on firms’ borrowing cost has also been studied in sectors other than the banking industry (see, for example, [Anginer and Warburton \(2014\)](#) for the auto industry). [Begenau and Stafford \(2018\)](#) propose that the reliance of banks on high leverage, presumably in the supply of liquidity, appears to generate costs of financial distress that are not offset with other benefits.

to our results is [Atkeson, d’Avernas, Eisfeldt, and Weill \(2018\)](#), who consider the extent to which TBTF affects the market-to-book ratios of banks, that is, the ratio of the market value of equity to the accounting value of equity. In principle, a post-crisis reduction in TBTF subsidies should lower market-to-book ratios. Indeed, the authors show that the equity-to-book ratio of publicly traded U.S. banks was above two, on average, between 1996 and 2007, and declined to about one after the 2008 financial crisis.

[Sarin and Summers \(2016\)](#) and [Chousakos and Gorton \(2017\)](#) argue, however, that the post-Lehman drop in bank market-to-book ratios is due to a loss in bank franchise value or profitability. Like us, [Atkeson, d’Avernas, Eisfeldt, and Weill \(2018\)](#) find that a substantial reduction in bank equity market values is instead associated with the decline of TBTF. They estimate⁶ that about 31% of their composite-bank market-to-book ratio was lost in the post-crisis period from a decline in government guarantees. This estimate is broadly consistent with our estimate of a 29% reduction in the market value of equity associated with the post-Lehman decline in bailout probabilities. (While these two estimated numbers are quite similar, the respective metrics are different.) To accomplish this, they construct a detailed dynamic model of the balance sheet and income statement of a single hypothetical composite U.S. bank, based on data for the aggregate U.S. banking sector, which consists of over 4,000 banks.⁷ They do not distinguish systemically important banks from other banks.

[Haldane \(2010\)](#) uses a ratings-based approach. He estimates the reduction in TBTF subsidies associated with the post-Lehman reduction in sovereign uplifts of the credit ratings of systemically important banks. Roughly speaking, for a given bank, [Haldane \(2010\)](#) assumes that the savings in its wholesale debt financing rates associated with TBTF can be estimated as the difference in average corporate bond yields associated with ratings that include and do not include the sovereign uplifts, respectively. Thus, smaller post-Lehman sovereign ratings uplifts automatically implies smaller estimated TBTF debt subsidies.⁸

⁶[Atkeson, d’Avernas, Eisfeldt, and Weill \(2018\)](#) estimate that the pre-crisis contribution of government guarantees to the market-to-book ratio was 0.91 and that the post-crisis contribution was about half of 1.19, for a reduction of about $0.91 - 0.60 = 0.31$.

⁷The number of U.S. banks is reported by the FDIC. To estimate the market value of equity of their modeled composite bank, [Atkeson, d’Avernas, Eisfeldt, and Weill \(2018\)](#) make the simplifying assumption that there are only two possible Markov states in each time period, and that the bank chooses to default in one of these, the crisis state.

⁸[Ueda and Weder di Mauro \(2013\)](#) provide estimates of the value of the subsidy to systemically important financial institutions (SIFIs) in terms of their credit ratings. They report that a one-unit increase in government support for banks in advanced economies has an impact equivalent to 0.55–0.90 notches on the overall long-term credit rating at the end of 2007. This effect increased to 0.80–1.23 notches by the end of 2009. [Rime \(2005\)](#) also examines the possible effects

Acharya, Anginer, and Warburton (2016) conduct an event study of the impact on G-SIB credit spreads of the passage of U.S. G-SIB failure resolution legislation, Title II of the Dodd-Frank Act. They find that there was no significant impact on G-SIB CDS rates within 60 days of the passage of Dodd-Frank. They also find that between 1990 and 2012, the bond credit spreads of the largest financial institutions are insensitive to risk, and that this is not the case for smaller financial institutions or for non-financial firms.

3. A High-Level Outline of Our Identification Strategy

At a very high level, our empirical strategy is to exploit the idea that a given corporate credit spread S can be approximated as the product of the annualized risk-neutral probability p of insolvency and the risk-neutral expected fractional loss ℓ in the event of insolvency. For a firm subject to bailout with risk-neutral probability π , the expected loss given insolvency is $\ell = (1 - \pi)L$, where L is the risk-neutral expected loss given insolvency with no bailout. The simple relationship $S = p(1 - \pi)L$ implies the log-linear model

$$\log \frac{S}{1 - \pi} = \log p + \log L. \quad (1)$$

We know from the results of Berndt, Douglas, Duffie, and Ferguson (2018) that the majority of the empirical variation in $\log p$ is explained by measured distance to default and time fixed effects, at least for a large sample of U.S. public firms. In the next section, we develop a theoretical framework that identifies channels through which bailout may impact a bank's measured distance to default. For the purposes of this high-level overview, we merely let $d_{it}(\pi_{it})$ denote the distance to default of firm i at date t corresponding to some assumed bailout probability π_{it} , given this firm's observed balance sheet at time t and other relevant data. We proxy for variation in $\log L$ using the vector of variables $l_{it}(\pi_{it})$.

So, based on the conceptual framework (1), we estimate bailout probabilities by fitting a model of the form

$$\log \frac{S_{it}}{1 - \pi_{it}} = \beta_0 + \beta_d d_{it}(\pi_{it}) + l_{it}(\pi_{it})\beta_l' + \omega_{K(i),t} + \varepsilon_{it}, \quad (2)$$

of TBTF expectations on issuer ratings and finds that proxies of the TBTF status of a bank have a significant, positive impact on bank issuer ratings.

where $\{\beta_0, \beta_d, \beta_l\}$ is the set of coefficients to be estimated, the term ω_{it} includes various fixed effects associated with the sector $K(i)$ of firm i and the date t , and ε_{it} is an unexplained residual. For parsimony, our fixed effects consist of (i) time fixed effects, (ii) sectoral fixed effects, and (iii) a sectoral fixed effect specific to big banks that is interacted with the time fixed effects. These fixed effects allow for variations over time and across sector in default risk premia (which affects the relationship between credit spreads, distance to default and loss given liquidation), and for deviations from the pricing relationship $S = p(1 - \pi)L$ due to market frictions.

For bailout probabilities, we take π_{it} to be zero if firm i is not a big bank. For a G-SIB bank i , we take $\pi_{it} = \pi_{\text{pre}}^G$ for all pre-Lehman dates and take $\pi_{it} = \pi_{\text{post}}^G$ for all post-Lehman dates, for two parameters, π_{pre}^G and π_{post}^G , to be estimated. We treat D-SIBs similarly, with respective bailout probabilities π_{pre}^D and π_{post}^D .

Within the underlying conceptual framework (2), it is challenging to identify the average level over time of the bailout probability π_{it} for big banks. For example, a higher sector-fixed effect due to higher default risk premia for big banks is observationally equivalent to a smaller average bailout probability π_{it} .⁹ We can, however, use the observable post-Lehman change in the relationship between credit spreads S , distance to default and loss given liquidation, caused by a regime shift in bailout expectations, to identify the *relative* change from $1 - \pi_{\text{pre}}^G$ to $1 - \pi_{\text{post}}^G$. Thus, as discussed in the introduction, we estimate a schedule of data-consistent pairs $(\pi_{\text{pre}}^G, \pi_{\text{post}}^G)$. For example, this schedule includes the pair (0.63, 0.2) and also the pair (0.67, 0.3). These two pairs of bailout probabilities, among other pairs, fit the data equally well. Schedules of estimated pairs of pre-Lehman and post-Lehman bailout probabilities for G-SIBs and D-SIBs, respectively, are shown in Section 7 and Appendix A.

For a given pair of pre-Lehman and post-Lehman big-bank bailout probabilities, the linear regression coefficients in (2) are chosen for the best fit of (2). We allow for heteroskedasticity and correlation in the residuals in a manner detailed in Section 7. Data-consistent pairs $(\pi_{\text{pre}}^G, \pi_{\text{post}}^G)$ are identified by forcing the degree to which the relationship between $\log[S_{it}/(1 - \pi_{it})]$, $d_{it}(\pi_{it})$ and $l_{it}(\pi_{it})$ for G-SIBs differs from that for non-banks to be the same in the pre-Lehman and the post-Lehman period, and likewise for D-SIBs.

Our numerical search for the best fitting parameters makes iterative use of the explicit solution

⁹This statement assumes that $\log(1 - \pi) + \beta_d d(\pi) + l(\pi)\beta_l'$ is a decreasing function of π , which is the case in our applications.

of the theoretical model presented in the next section, which captures the dependence of distance to default $d_{it}(\pi_{it})$ and the LGL proxies $l_{it}(\pi_{it})$ on bailout probability and on various observable balance-sheet variables.

4. Valuation of Bank Equity and Debt with Bailout Subsidies

This section presents a simple model of the valuation of a bank’s debt and equity, capturing the effects of a given probability of government bailout at insolvency. The model captures the impact of bailout on solvency, credit spreads, equity market value, and the component of equity market value associated with bailout subsidized debt.

We consider a bank whose assets in place, V_t , at any time t before default, satisfy the stochastic differential equation

$$dV_t = V_t(r - k) dt + V_t\sigma dZ_t, \quad (3)$$

where Z is a standard Brownian motion under some risk-neutral probability measure,¹⁰ r is the risk-free interest rate, σ is the asset volatility, and k is the total cash revenue rate as a proportion of assets in place. At any default time τ , if the bank is liquidated in a bankruptcy process, distress costs cause the value of assets in place to drop from $V_{\tau-}$ to $V_{\tau} = \alpha V_{\tau-}$, for some recovery coefficient $\alpha \in (0, 1)$.

Our most basic model of the bank has two layers of debt: deposits and bonds. Details are provided in Appendix B. Appendix C considers an extension with three layers of debt: deposits, senior bonds, and junior bonds. With data bearing separately on the pricing of junior and senior debt, it would be possible in principle to estimate the likelihood of a bail-in failure resolution, which focuses insolvency losses more heavily on junior debt. This is beyond the scope of our current study, which lumps bail-in together with other insolvency procedures that generate losses for unsecured wholesale debt claims.

Deposits are of constant total size D and pay interest dividends at some constant rate d . Deposits are guaranteed, at no cost to the bank, by the government. Appendix B.3 extends the model and

¹⁰We fix a probability space $(\Omega, \mathcal{F}, \mathbb{P})$ and a filtration $\{\mathcal{F}_t : t \geq 0\}$ of sub- σ -algebras of \mathcal{F} satisfying the usual conditions. For details see, for example, Protter (2005). All of our probabilistic statements in this section are relative to a probability measure \mathbb{Q} , equivalent to \mathbb{P} , under which the market value at time t of a claim to any increasing adapted cumulative cash-flow process C is $E^{\mathbb{Q}}\left(\int_t^{\infty} e^{-r(u-t)} dC_u \mid \mathcal{F}_t\right)$, where r is the given short rate. The process Z is a standard Brownian motion with respect to the probability space $(\Omega, \mathcal{F}, \mathbb{Q})$ and filtration $\{\mathcal{F}_t : t \geq 0\}$. The probability measure \mathbb{Q} is called a “risk-neutral” measure.

its explicit solution so as to accommodate FDIC deposit insurance premia and the additional FDIC assessments that were introduced after the financial crisis for the liabilities of “large and complex financial institutions.” The effective level of the post-crisis FDIC assessment rates for each bank is difficult to determine and not directly observable. The post-crisis assessment rates are believed by some commenters (Whalen, 2011; Pozsar, 2016) to have ranged from ten basis points, on average, when first introduced, then declining to around five basis points at the end of our sample period. For simplicity, we do not include the effects of FDIC insurance assessments in our empirical estimation.

Because of imperfect competition in the deposit market, deposit rates are in practice much lower than wholesale risk-free rates, and rise very sluggishly after risk-free rates rise (Driscoll and Judson, 2013; Drechsler, Savov, and Schnabl, 2017). In our empirical implementation of the model, this is reflected in our data inputs for d and r .

The remaining class of debt consists of bonds of constant total principle P , with maturities that are exponentially distributed.¹¹ That is, bonds mature at some aggregate proportional rate $m > 0$, so that the fraction of the original bond principle that remains outstanding at any time t is e^{-mt} . This implies that the average bond maturity is $1/m$. The bonds have some coupon rate of c per unit of principle, for a total coupon payment rate of cP on all outstanding bonds. When any existing bond matures at time t , the same principle amount of debt is issued at its current market value, which could be at a premium or discount to par depending on V_t . The newly issued bonds have the original coupon rate c . The original exponential maturity distribution is always maintained. Interest payments are tax deductible at the corporate tax rate κ . This is the model of Leland (1994a), except that we have two classes of debt, insured deposits and bonds, and that we allow for a government bailout at default.

The bank’s current shareholders choose a stopping time τ at which they will no longer service the bank’s debt. That is, at time τ the current equity owners default, and stop participating in any cash flows, permanently. We focus on an equilibrium default time of the form $\tau = \inf\{t : V_t \leq V^*\}$, for some constant asset default boundary V^* . The empirical results presented in the paper are for the case

¹¹A constant exponential maturity structure can be achieved as follows, among other assumptions. There could be initially a continuum (non-atomic measure space) of different bonds with aggregate principle P . (The principle of each bond is “infinitesimal.”) The maturity date of each bond is random and exponentially distributed with parameter m . The maturity dates are pairwise independent. The measurability conditions of Sun (2006) can be used to support an application of the exact law of large numbers, under which the cross-sectional distribution of maturity dates is equal to the same exponential distribution. Each time a bond matures, it pays the investor its principle and is replaced by issuing at market value a new bond of the same principle whose maturity is exponentially distributed with the same parameter m , again independently of all else.

where V^* , and thus τ , are chosen by shareholders to maximize the market value of their equity claim.¹² In Appendix D, we obtain similar results for an exogenous specification of V^* that is consistent with historical liquidation bond recovery rates. The market value of the bank’s equity converges to zero as t approaches τ .

At the default time τ , the bank does not necessarily go into an insolvency process, causing distress costs and lack of bond payment. The bank could instead be “bailed out” by the government. Bailout is not predictable¹³ and occurs with a given risk-neutral probability π . That is, the conditional probability at any time t of bailout at the next default is always equal to the unconditional bailout probability π . Although the original equity owners default on their debt obligations at time τ , in the event of a bailout the government injects new capital, thus avoiding any distress costs to assets. The government becomes the new equity owner. The debt continues to be serviced by the bank as originally contracted.

This is roughly what happened¹⁴ with many of the bailouts of large European banks, with the nationalizations of, or dominant majority government equity positions taken in, Royal Bank of Scotland, Northern Rock, KBC, Fortis, Dexia, ABN Amro, SNS Reaal, Hypo Real Estate Bank, Anglo Irish Bank, and Bank of Ireland. In practice, a government might instead partially nationalize a bank before the market value of its equity reaches zero, as with some other European bailouts (including those of Lloyds, Commerzbank, UBS, ING, Bankia, Alpha Bank, Eurobank, and Piraeus Bank) and with the U.S. Troubled Asset Relief Program (TARP). For simplicity, we avoid the modeling of partial nationalization.

Immediately after a bailout, the government may sell its equity stake on a competitive market and the bank continues to operate, following the same policy, until at least the next such default time, and so on. The government’s bailout capital injection is, for simplicity, modeled as a purchase of some quantity $\widehat{V} - V^*$ of additional assets. The quantity of additional assets is precisely enough to bring the total market value of the bonds up to some stipulated value B . The net government bailout subsidy is

¹²When asset default boundary V^* is determined endogenously, our solution concept is the equilibrium default timing model of [Décamps and Villeneuve \(2014\)](#), by which debt is issued at each time at a competitive market price that is consistent with correct investor conjectures of the default-time policy τ .

¹³The event of bankruptcy or bailout is revealed precisely at time τ . That is, we can define the commonly observed information filtration $\{\mathcal{F}_t : t \geq 0\}$ by letting $\mathcal{F}_t = \sigma(\{z_s : s \leq t\} \cup \{B_1, \dots, B_n\})$, where B_1, B_2, \dots is a risk-neutrally independent sequence of Bernoulli trials corresponding to successive bailouts of the bank (one if bailout, zero otherwise), and B_n is the last such trial that has occurred by time t . The trials are revealed each time that V_t reaches V^* .

¹⁴See “What has happened to more than 30 bailed-out European banks,” Reuters, August 21, 2015, at <https://uk.reuters.com/article/europe-banks-bailouts/factbox-what-has-happened-to-more-than-30-bailed-out-european-banks-idUKL5N10W0XJ20150821>.

thus $\widehat{V} - V^* - H(\widehat{V})$, where $H(x)$ is the market value of equity for a bank with initial assets in place of x and with the original liability structure (D, d, P, c) .

In the event of no bailout at a default time, the bank is permanently liquidated. At liquidation, the deposits are redeemed in full, using if necessary the government's deposit guarantee, and the bond creditors receive any remaining liquidation value of assets, pro rata by principal amount. That is, per unit of the face value of their debt claims, the depositors receive one unit of value and the bond holders receive $(\alpha V^* - D)^+/P$. Here x^+ denotes the positive part of x , meaning $x^+ = \max(0, x)$. Government default insurance pays $(D - \alpha V^*)^+$.

This simple model is time-homogeneous with Markov state variable V_t . The vector of primitive model parameters is

$$\Theta = (c, P, B, m, d, D, r, k, \sigma, \alpha, \kappa, \pi).$$

We will only consider parameters for which there is non-trivial default risk for creditors, meaning that the recovery value $(\alpha V^* - D)^+$ of bonds in bankruptcy does not exceed the recovery B at bailout, which in turn does not exceed the default-risk-free market value of the bonds, $P(c + m)/(r + m)$. Because we have an explicit solution for V^* in terms of primitive model parameters, this non-degeneracy condition is an explicit condition on these model parameters. The model can also be solved explicitly in the degenerate case, which we avoid only for simplicity of notation.

We now turn to the market valuation of various relevant contingent claims, and a calculation of the optimal default boundary V^* and the bailout recapitalized asset level \widehat{V} . The detailed calculations and explicit solutions are found in Appendix B. Here, we just provide an overview of the solution method.

At first taking the default and recapitalization boundaries V^* and \widehat{V} as given, we compute the total market value of all cash flows available to the bank over the time period $[0, \infty)$, including the cash flows generated by the original assets in place, government tax shields, deposit insurance payments, additional rents or costs on the asset side, and bailout capital injections, net of distress costs, at all future bailouts. We then equate this total value of net available cash flows to the total market value of all of the positions held by claimants against the same net cash flows. These claimants are the original equity owners, the original depositors, the original bond holders, and the government as a contingent equity claimant at all future bailouts. From this equation, taking V^* as given, we can explicitly deduce

the market value of equity and the recapitalized asset level \widehat{V} , in terms of the default boundary V^* . Finally, using a “smooth pasting” condition for the market value of equity at the default boundary V^* , we solve explicitly for V^* .

For a simple illustrative set of parameters, Figure 1 plots the market values of a payment of \$1 at insolvency, $U(V_0)$, the market value of equity, $H(V_0)$, and the default and recapitalization asset boundaries, V^* and \widehat{V} , showing how these variables depend on the assumed bailout probability π . Naturally, the market values of equity is increasing in π . As π goes up, the endogenous asset default boundary V^* goes down, reflecting the incentive of shareholders to extend the period of time over which they enjoy higher subsidies in their debt financing costs, caused by higher expectations of creditor bailout at any future default. Figure A.1 in the appendix reveals that most of the total value of bailout subsidies are picked up by the initial equity owners, through the prospect of future and more heavily subsidized debt re-issuance prices that will be available if and when asset levels decline unexpectedly.

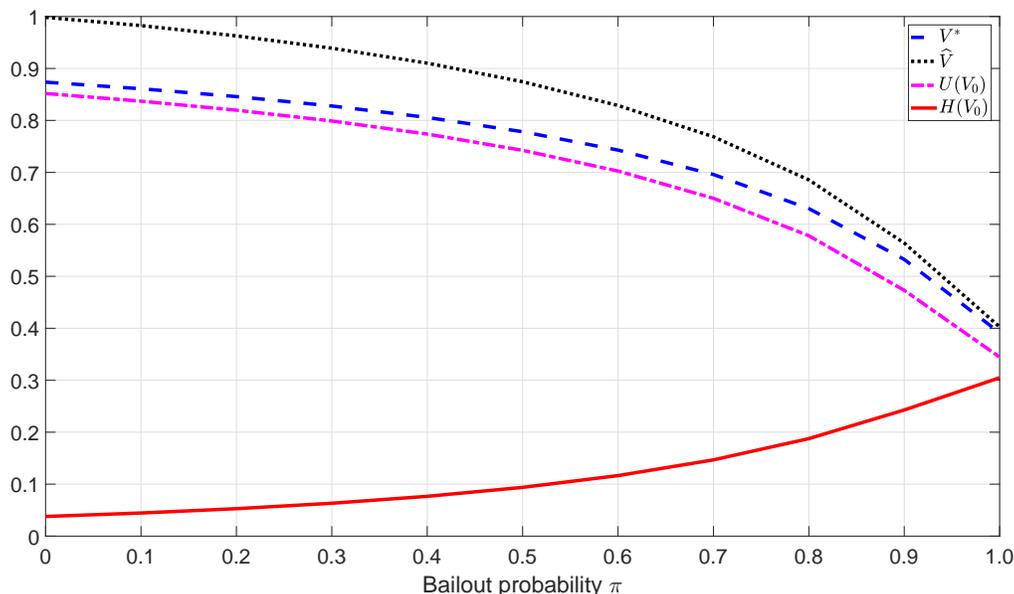


Figure 1: **Effect of bailout on theoretical claim valuations and endogenous asset boundaries.** The figure plots the explicit solutions of the model valuation $U(V_0)$ for a payment of \$1 at insolvency, $H(V_0)$ for equity, the asset default boundary V^* , and the bailout asset level \widehat{V} , as functions of the bailout probability π . There are no deposits. The other underlying primitive model parameters are $V_0 = 1$, $\sigma = 0.1317$, $\alpha = 0.5$, $r = 0.0230$, $k = 0.0221$, $\phi = -0.0136$, $\kappa = 0.35$, $m = 0.4233$, $c = 0.0680$, and $B = P = 0.5137$. These parameter choices are consistent with the calibrated model for Goldman Sachs in mid-September 2008, scaled to $V_0 = 1$.

In prior work that incorporates government bailouts into dynamic models of endogenous default

and bond pricing,¹⁵ Chen, Glasserman, Nouri, and Pelger (2017) and Albul, Jaffee, and Tchisty (2010) assume that the government simply guarantees (with probability one) the debt principal at default.¹⁶ Our empirical objectives obviously require a model that allows variation in the bailout probability. Taking a “macro” approach, Gandhi, Lustig, and Plazzi (2016) assume that the government absorbs the aggregate losses of the entire financial sector, above an assumed cap, that could be caused by a “rare disaster.”

5. Data and Descriptive Statistics

The focus of our analysis is systemically important banks. The Financial Stability Board (FSB), in consultation with the Basel Committee on Banking Supervision (BCBS) and national authorities, identifies “global systemically important banks” (G-SIBs). Our analysis is based on the list of G-SIBs published in November 2017.¹⁷ Eight U.S. bank holding companies are identified as G-SIBs: Bank of New York Mellon, Bank of America, Citigroup, Goldman Sachs, JPMorgan Chase, Morgan Stanley, State Street and Wells Fargo.

As a separate category of big banks, we also treat those U.S. banks, beyond G-SIBs, that are sufficiently systemic to require stress tests under the Fed’s Comprehensive Capital Analysis and Review (CCAR) and Dodd-Frank Act stress test (DFAST).¹⁸ As of 2018, there are eighteen such firms: Ally Financial, American Express, BB&T, Capital One Financial Corp, CIT Group, Citizens Financial Group, Comerica, Discover Financial Services, Fifth Third Bancorp, Huntington Bancshares, KeyCorp, M&T Bank, Northern Trust, PNC Financial Services Group, Regions Financial, Suntrust Banks, U.S. Bancorp, and Zions Bancorporation, which we label “domestic systemically important banks” (D-SIBs).

In addition to G-SIBs and D-SIBs, we also collect data on all other public U.S. firms that can be matched unambiguously across the Markit CDS, Compustat and CRSP databases. Markit credit

¹⁵Other work applying variations and extensions of the Leland (1994b) framework to financial firms includes Auh and Sundaresan (2018), Diamond and He (2014), Harding, Liang, and Ross (2013), He and Xiong (2012), and Sundaresan and Wang (2014).

¹⁶Albul, Jaffee, and Tchisty (2010) have perpetual debt, so, once the debt is issued at time zero, the government guarantee of debt has no effect on the default boundary or on the market valuation of equity. With Chen, Glasserman, Nouri, and Pelger (2017), on the other hand, there is issuance of new debt over time as in our model. Chen, Glasserman, Nouri, and Pelger (2017) and Albul, Jaffee, and Tchisty (2010) allow for the risk of unexpected jumps in asset value. We have avoided jumps, merely, for simplicity.

¹⁷This list is based on the use of end-2016 data and on the updated assessment methodology published by the BCBS in July 2013 (Financial Stability Board, 2017; Bank of International Settlement, 2013).

¹⁸The FED report is available at www.federalreserve.gov/newsevents/pressreleases/bcreg20180201a.htm, and was accessed on June 25, 2018.

default swap (CDS) rate observations are “at-market,” meaning that they represent bids or offers of the default-swap rates at which a buyer or seller of protection is proposing to enter into new default swap contracts without an up-front payment. The at-market CDS rate is in theory that for which the net market value of the contract is zero, assuming no upfront and zero dealer margins. The rates provided by Markit are composite CDS quotes, in that they are computed based on bid and ask quotes obtained from three or more anonymous CDS dealers.

We use CDS data based on a contractual definition of default known as “no restructuring.” This contractual definition allows for CDS protection against losses in the event of a bankruptcy or a material failure by the obligor to make payments on its debt. Our CDS data apply to senior unsecured debt instruments, and are available for maturity horizons from one to ten years. For banks, our data cover CDS for holding-company bonds.

We only use CDS quotes for which Markit rates the data quality of the quotes as “BB” or higher. If a quote-quality rating is not available, we require a composite level of “CcyGrp,” “DocAdj” or “Entity Tier.”¹⁹ Although Markit CDS data go back as far as 2001, after cleaning the data we find few 2001 observations. We therefore restrict our sample to the period from 2002 to 2017. Lastly, we exclude firms with less than one year of CDS data.

Accounting data are available from quarterly Compustat files. Items downloaded include book assets, long-term debt, short-term debt, cash dividends, and interest expenses. Whenever quarterly data are missing, we use annual reports to augment the data. To avoid a forward-looking bias, on any given date we use the accounting data from the last available quarterly report. For large banks—meaning G-SIBs and D-SIBs—we use the Compustat Banks database as well as 10-Q and 10-K SEC disclosure filings to capture other data.²⁰

For each large bank, we calculate a daily time series of notional-weighted average bond maturities using the maturity information provided in 10-K filings. For all other firms, we approximate by treating the maturities of short-term and long-term bonds as though equal to one year and five years, respectively, then using the reported notional amounts of short-term and long-term debt to compute

¹⁹To meet these data quality requirements, all but four D-SIBs would have to be excluded from the sample. To obtain more robust results, we do not impose these restrictions for D-SIBs.

²⁰We have verified that the information contained in the 10-K filings of large banks closely matches that available through Compustat, especially for book assets, long-term debt, and deposits. We use Compustat as our main data source because of the consistency that it offers in terms of measuring short-term debt. The 10-K definition of short-term debt, by comparison, is inconsistent across banks and time.

notional-weighted bond maturities.

Equity market data, including the number of shares outstanding and price per share, are obtained from CRSP.

The final sample contains 637 unique firms—as identified by their CRSP “permco” number—from ten industry sectors. It includes six of the eight G-SIBs (Bank of America, Citigroup, Goldman Sachs, JPMorgan Chase, Morgan Stanley and Wells Fargo) and nine D-SIBs (American Express, BB&T Corp, Capital One Financial Corp, CIT Group, Discover Financial Services, KeyCorp, PNC Financial Services Group, Suntrust Banks Inc and U.S. Bancorp). Other big banks are excluded because of the lack of data or insufficient data quality.

The range of credit qualities of the firms in our data may be judged from Table 1, which categorizes firms according to their median Moody’s rating over the sample period. The table shows, for each credit rating, the number of firms in our study with that median rating. As the table indicates, firms in the sample tend to be of medium credit quality. Across industry groups, ratings tend to be higher for financial, healthcare, and technology firms, and lower for telecommunication services firms.

Table 1: Distribution of firms across sectors and by credit quality. The table reports the distribution of firms across sectors and by median Moody’s senior unsecured issuer ratings. The data include 637 public U.S. firms, over the period 2002–2017.

	Aaa	Aa	A	Baa	Ba	B	Caa	Ca-C	NR	All
Basic Materials	0	0	7	20	15	6	1	0	1	50
Consumer Goods	0	3	15	41	24	17	6	0	3	109
Consumer Services	0	2	10	36	20	18	11	2	5	104
Energy	0	1	4	32	8	8	1	0	3	57
Financials	1	3	26	42	9	3	1	0	6	91
Healthcare	0	1	6	15	9	4	1	0	3	39
Industrials	1	2	14	24	16	14	4	0	6	81
Technology	0	0	4	10	3	6	0	0	3	26
Telecommunications Services	0	0	4	7	4	3	3	0	1	22
Utilities	0	0	5	35	6	8	1	0	3	58
All	2	12	95	262	114	87	29	2	34	637

Appendix Figure A.2 shows time series of median five-year CDS, for G-SIBs, D-SIBs and all other firms. Median CDS rates are substantially higher following WorldCom’s default in July 2002, during the 2008-09 financial crisis, and during the latter half of 2011, when there were severe concerns about European peripheral sovereign debt and faltering negotiations over the U.S. government debt ceiling. The increase in CDS rates in late 2011 was particularly pronounced for G-SIBs.

Table 2 reports summary statistics for certain key accounting variables for G-SIBs, D-SIBs and other firms. We consider two sub-periods: the “pre-Lehman” period from January 1, 2002 to the date of Lehman Brothers bankruptcy on September 15, 2008, and the “post-Lehman” period from September 16, 2008 to December 31, 2017. As shown, large banks tend to have much higher accounting leverage than other firms, yet tend to have much lower market rates for CDS protection, especially before the financial crisis. Our leverage data include only conventional debt liabilities, and not the liabilities associated with over-the-counter derivatives, repurchase agreements, and other qualified financial contracts (QFCs). Several of the G-SIBs, Goldman Sachs, Morgan Stanley, J.P. Morgan, Citigroup, and Bank of America have substantial amounts of QFCs.

Table 2: **Accounting measures** This table reports averages statistics for book assets (BVA), book debt (BVD), short-term debt (STD), long-term debt (LTD), deposits (Dpst), market capitalization (MC), cash dividends (CD), and interest expense (IE), in billions of U.S. dollars. Book leverage (Lev) is the ratio of the sum of book debt and deposits to book assets. (Unconventional liabilities, such as those associated with over-the-counter derivatives, are not included.) Sample-average five-year CDS rates are reported in basis points. Moody’s ratings (Rtg) are coarse letter ratings. Notional-weighted average bond maturities (Mat) are shown in years. The pre-Lehman period (Pre) is January 1, 2002 to September 15, 2008. The post-Lehman period (Post) is September 16, 2008 to December 31, 2017.

	BVA	BVD	STD	LTD	Dpst	MC	CD	IE	Lev	CDS	Rtg	Mat	Firms
G-SIBs													
Pre	949	377	251	126	317	119	4.01	20.44	0.73	38	Aa	2.48	6
Post	1599	489	255	234	690	133	3.65	12.85	0.70	126	A	2.98	6
D-SIBs													
Pre	118	34	11	23	56	28	0.72	2.09	0.76	75	A	3.68	9
Post	191	35	8	27	117	31	0.71	1.56	0.74	158	Baa	4.14	9
Other firms													
Pre	31	15	7	8	0	16	0.34	0.16	0.51	143	Baa	3.07	561
Post	44	17	6	10	0	25	0.74	0.13	0.53	227	Baa	3.31	378

In our current empirical estimation, we have not captured the effects of some significant mergers. During our sample period, J.P. Morgan acquired Bank One (2004) and Bear Stearns (2008), Bank of America acquired Merrill Lynch (2008), and Wells Fargo acquired Wachovia (2008). These mergers have relatively little impact on our post-Lehman data set, so should not have a large impact on our estimates of post-Lehman versus pre-Lehman bailout probabilities. Nevertheless, our failure to capture the effects of these mergers may somewhat bias our results, and add noise. In a future revision, we plan to build pro-forma merged datasets for these firms.²¹

²¹Accounting data for debt liabilities and market values of equity can simply be added together. We can construct pro-forma estimates of CDS rates, as weighted averages of the pre-merger CDS rates of the constituent firms, weighting by principle amounts P of non-deposit debt.

6. Model Parameters

As explained in Section 3, we use a panel-regression approach to estimate bailout probabilities from corporate credit spreads, estimated distance to default and loss given liquidation (LGL) proxies. In order to calculate the distance to default and LGL proxies corresponding to a given assumed bailout probability π , we need to select inputs for all of the remaining theoretical model parameters $(c, P, D, B, m, r, r^d, k, \phi, \sigma, \alpha, \kappa)$.

In practice, any model is mis-specified to some extent. We face a tension between the restrictive theoretical assumption that these primitive model parameters are treated by agents in the model as though constant over time and the practical need in our non-stationary setting to allow some of these parameters to vary across observation dates for the purpose of estimating distance to default and bailout probabilities. This section is a summary of how each of the model parameters other than π , which is a fitted variable, is chosen as an input to the theoretical model for the purpose of computing distance to default and LGL proxies. Further details on parameter selection are provided in Appendix E.

The proportional asset recovery α at bankruptcy and the corporate tax rate κ are fixed for the entire panel study at 50% and of 35%, respectively, common to all firms. For bank i on date t , the quantity D_{it} of deposits and the average deposit interest rate r_{it}^d are obtained from accounting disclosure for the corresponding quarter. Of course, for non-banks, there are no deposits. The total principle P_{it} of non-deposit debt (both short-term and long-term) and the notional-weighted average bond maturity $1/m_{it}$ are also obtained, quarter by quarter, from public firm disclosure. The risk-free rate r_{it} that we use as a theoretical model input for firm i on date t is the constant-maturity treasury (CMT) rate interpolated to the average maturity $1/m_{it}$ of the firm’s non-deposit debt.²²

For each firm, on each individual date, the modeled non-deposit debt coupon rate c_{it} is taken to be the sum of (i) the risk-free rate r_{it} and (ii) an estimate of the bond yield spread for that firm, that date. The estimated bond yield spread is the firm’s at-market credit default swap (CDS) rate, as linearly interpolated from the data to the firm’s average non-deposit debt maturity $1/m_{it}$, net of the CDS-bond basis.²³

For large banks, the market value B_{it} of non-deposit debt immediately after a bailout is set equal to

²²We download the constant-maturity treasury rate from the website of Federal Reserve Bank of St. Louis, <https://fred.stlouisfed.org/categories/115>.

²³If at time t firm i has investment-grade (IG) status, we subtract the Markit IG CDX basis. If the firm has high-yield (HY) status, we subtract the Markit HY CDX basis.

par, $B_{it} = P_{it}$. In Appendix D, we consider an alternative specification where B_{it} is chosen to achieve a given post-bailout bond yield spread s , by setting $B_{it} = P_{it}(c_{it} + m_{it})/(r_{it} + s + m_{it})$. This is based on the idea that a government bailout would target a given level of creditworthiness of a large bank as judged in wholesale credit markets. We set $s = 100$ basis points (bps), which is reasonable but still somewhat arbitrary. Results are qualitatively similar.

The volatility parameter σ_{it} of a given firm i on a given date t is assumed to be constant across dates within the pre-Lehman period, and in the post-Lehman period to be constant at a different level. Within each of these two periods, σ_{it} is taken to be the annualized sample standard deviation of first differences of the logarithm of the model-implied asset levels V_{it} of firm i within that sub-period. The model-implied asset level V_{it} in turn depends on the assumed volatility σ_{it} , so this calibration involves an iterative search for σ_{it} .

The firm payout rate k_{it} is set at some multiple ρ_{it} of the risk-free rate r_{it} . We impose an over-identifying restriction by assuming that ρ_{it} is constant within each of the pre-Lehman and post-Lehman periods. The constant level of ρ_{it} within a given period (whether pre-Lehman or post-Lehman) is chosen to match the within-period average of $k_{it} + \phi_{it}$ to the within-period average ratio²⁴ of (i) the sum of cash dividends and interest expenses to (ii) the model-implied assets V_{it} . The model-implied asset level V_{it} depends in turn on ρ_{it} , so we conduct a joint iterative search for σ_{it} and ρ_{it} for consistency of the resulting time series of V_{it} with our fitting equations for σ_{it} and ρ_{it} . Further details are provided in Appendix E. In Appendix D, we present our findings for an alternative measure of interest expense and show that the results are quantitatively similar.

In the main part of the paper, we present results for the case where shareholders choose the optimal insolvency boundary level of assets. In Appendix D, we verify that results remain qualitatively similar for an expected recovery at insolvency given no-bailout of 40% of notional. This estimate of the liquidation bond recovery rate $(\alpha V^* - D)^+/P$ falls between two recovery rate measures reported by [Moody's Investor Services \(2018\)](#). Moody's reports a historical issuer-average ultimate recovery rate for senior unsecured bonds, for 1987–2017, of 47.9%, and an average recovery rate measured by

²⁴That is, letting p denote a given sub-period, whether pre-Lehman or post-Lehman, ρ_{it} is constant across t in that sub-period, and set so that $\sum_{t \in p} (k_{it} + \phi_{it}) - C_{it}/V_{it} = 0$, where $k_{it} = \rho_{it} r_{it}$, and C_{it} is the sum of cash dividends and interest expense for firm i on date t , and V_{it} is the firm's model-implied asset level on date t . We restrict ρ to the interval $(0, 1)$, to ensure that $0 < k < r$. In cases where ρ converges towards the boundary value of one, we require $\sum_{t \in p} (k_{it} + \phi_{it})$ to be as close as possible to $\sum_{t \in p} C_{it}/V_{it}$.

the market prices of bonds immediately after default, over 1983–2017, of 37.7%.

7. Model Estimation and Empirical Results

The distance to default d of a generic firm in our theoretical setting is defined by

$$d_t = \frac{\log(V_t) - \log(V^*)}{\sigma}, \quad (4)$$

where V_t is current assets in place, V^* is the endogenous asset default boundary, and σ is the volatility of assets. Variation in loss given liquidation is captured by two proxies,

$$l_t^1 = \frac{\log(P_t + D_t) - \log(V^*)}{\sigma} \quad \text{and} \quad l_t^2 = \frac{\log(P_t) - \log(P_t + D_t)}{\sigma}, \quad (5)$$

where l_t^1 is the risk-adjusted distance between book liabilities and the default threshold and l_t^2 relates to the liability structure of the firm.²⁵ The specification of the LGL proxies in (5) allows us to compute the solvency ratio d_t^s of the firm as

$$d_t^s = d_t - l_t^1 = \frac{\log(V_t) - \log(P_t + D_t)}{\sigma}. \quad (6)$$

The solvency ratio measures the risk-adjusted distance between current assets and book liabilities.

Following the strategy previewed in Section 3, we now describe more precisely our estimation of pre-Lehman bailout probabilities, π_{pre}^G for G-SIBs and π_{pre}^D for D-SIBs, for various assumed values of the corresponding post-Lehman bailout probabilities π_{post}^G and π_{post}^D .

Our basic empirical model is

$$\begin{aligned} \log \frac{S_{it}}{1 - \pi_{it}} = & \beta_0 + \beta_d d_{it}(\pi_{it}) + \beta_{l,1} l_{it}^1(\pi_{it}) + \beta_{l,2} l_{it}^2(\pi_{it}) + \sum_{\text{sector } j} \delta_j D_j(i) + \sum_{\text{month } m} \delta_m D_m(t) \\ & + \delta_{\text{post}}^G D_{\text{post}}(t) D^G(i) + \delta_{\text{post}}^D D_m(t) D^D(i) + \varepsilon_{it}, \end{aligned} \quad (7)$$

²⁵Expected fractional loss given liquidation is given by

$$L_t = \frac{P_t - (\alpha V^* - D_t)^+}{P_t} = \left(1 - \max \left\{ \alpha \frac{V^*}{P_t + D_t}, \frac{D_t}{P_t + D_t} \right\} \right) \frac{P_t + D_t}{P_t},$$

and thus is a function of l_t^1 and l_t^2 .

where S_{it} is the five-year CDS rate of firm i on date t ;

$$C_L = (\beta_0, \beta_d, \beta_{l,1}, \beta_{l,2}, \{\delta_j : j \in \text{sectors}\}, \{\delta_m : m \in \text{months}\}, \delta_{post}^G, \delta_{post}^D)$$

are the linear-model coefficients to be estimated; $D_j(i)$ is the indicator (0 or 1) for whether firm i is in a sector j ; $D_m(t)$ indicates whether date t is in month m ; $D_{post}(t)$ indicates whether date t is in the post-Lehman period; $D^G(i)$ indicates whether firm i is a G-SIB; $D^D(i)$ indicates whether firm i is a D-SIB; and ε_{it} is an uncertain residual.

We take π_{it} to be zero if firm i is not a large bank. If firm i is a G-SIB, we take π_{it} to be π_{pre}^G for any pre-Lehman date t and π_{post}^G for any post-Lehman date t . Likewise, if firm i is a D-SIB, we take π_{it} to be π_{pre}^D for any pre-Lehman date t and π_{post}^D for any post-Lehman date t . Holding π_{post}^G and π_{post}^D fixed as model parameters, we search for π_{pre}^G and π_{pre}^D with the property that the fitted version of (7) satisfies

$$\delta_{post}^G = \delta_{post}^D. \quad (8)$$

Our model thus allows for temporal variation in default risk premia over months, and allows this variation to be different across G-SIBs, D-SIBs, and the non-bank sectors. However, Equation (8) forces the degree to which default risk premia for G-SIBs differ proportionately from default risk premia in the non-bank sector to be the same on average in the pre-Lehman and as in the post-Lehman period, and likewise for D-SIBs. This condition is achieved by choice of the bailout probabilities π_{pre}^G and π_{pre}^D .

For any given bailout probability parameters, we estimate the coefficients C_L of the linear model (7) as a standard panel regression. Figure 2 shows the resulting estimates for δ_{post}^G and δ_{post}^D , as a function of various pre-Lehman bailout probabilities π_{pre}^G and π_{pre}^D , for the special case in which the post-Lehman bailout probabilities for large banks are set to 0.2. As shown, the identifying constraints (8) are satisfied for $\pi_{pre}^G = 0.63$ and $\pi_{pre}^D = 0.46$. Alternatively, when the post-Lehman bailout probability is set to zero, the identifying constraints are satisfied for $\pi_{pre}^G = 0.54$ and $\pi_{pre}^D = 0.39$.

The resulting coefficient estimates are equivalent to those obtained when estimating all of the parameters $(C_L, \pi_{pre}^G, \pi_{pre}^D)$ of the non-linear model (7), subject to (8), by non-linear least squares. All of the coefficient estimates and standard errors are reported in Appendix Table A.1. The root mean

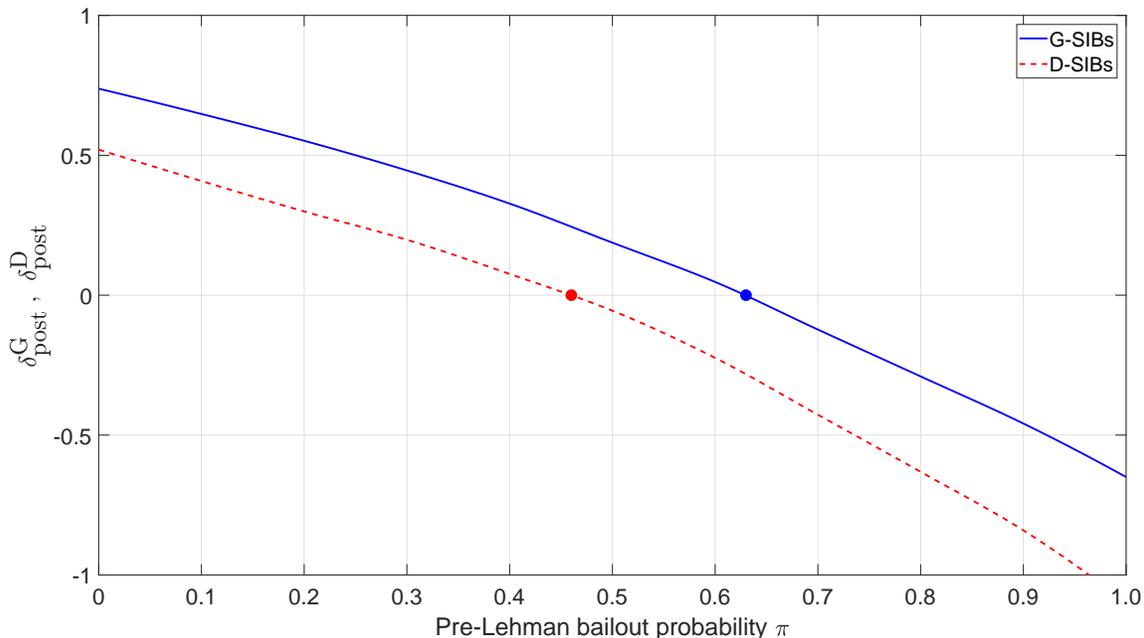


Figure 2: **Post-Lehman big-bank time fixed effects.** This figure shows the estimates for $\bar{\delta}_{\text{post}}^G$ and $\bar{\delta}_{\text{post}}^D$ in the panel regression (7), as a function of the pre-Lehman bailout probability π_{pre}^G . The post-Lehman bailout probabilities for large banks are set to 0.2. As shown, the identification condition (8) implies pre-Lehman bailout probability estimates of $\pi_{\text{pre}}^G = 0.63$ and $\pi_{\text{pre}}^D = 0.46$.

squared error (RMSE) for the fitted relationship is 0.46. While the CDS data are noisy in this sense, the relationship between the log CDS rate and the distance to default is highly significant. Variation in distance to default, sector and month fixed effects explain a sizable fraction—an R^2 of about 0.78—of variation in log CDS rates.

The solid blue line plotted in Figure 3 shows the fitted credit spread (five-year CDS rate) for G-SIB holding company senior unsecured bonds at a distance to default of 2.0. As shown, the cost of debt financing for G-SIBs at this fixed level of insolvency risk is on average much higher after the crisis. The dashed red line in Figure 3 indicates that some of this post-Lehman increase in debt financing costs, at a fixed solvency ratio, is caused by a general post-Lehman increase in market default risk premia. That is, the red line is the fitted G-SIB credit spread for the counterfactual case in which bailout probabilities did not go down after the crisis. The remainder of the post-Lehman increase in big-bank debt financing costs, the difference between the blue and red lines, reflects the impact of a substantial post-Lehman drop in the fitted bailout probability. As shown in the figure, at a fixed insolvency risk, post-crisis credit spreads were roughly double what they would have been had pre-Lehman bailout

probabilities been maintained in the post-crisis period.

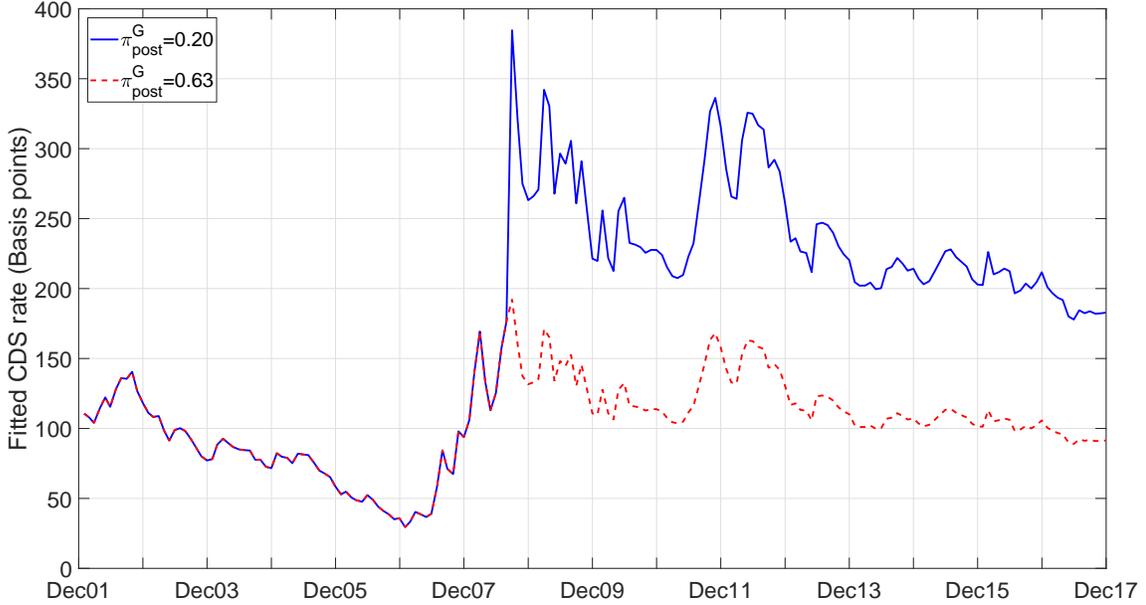


Figure 3: **Fitted CDS rates.** This figure shows fitted credit spreads (5-year CDS rates) for G-SIBs at a solvency ratio of two standard deviations, assuming the default threshold is equal to book liabilities and that there are no deposits. The solid blue line is based on the fitted pre-Lehman bail-out probability of $\pi_{pre}^G = 0.63$, and an assumed post-Lehman bail-out probability $\pi_{post}^G = 0.20$. The red line is based on the counterfactual assumption that there is no post-Lehman change in bail-out probability, taking $\pi_{post}^G = 0.63$.

Taylor and Williams (2009) point out that there was actually a significant increase in big-bank credit spreads about a year *before* the failure of Lehman. Their metric for big-bank credit spreads is the spread between three-month LIBOR and the three-month overnight index swap (OIS) swap rate, a proxy for close-to-risk-free borrowing rates. In August 2007, this three-month credit spread for large banks rose from about twenty basis points to about 80 basis points, most likely on news about heightened default risks in the U.S. residential mortgage market. In the context of our model and its motivating hypothesis, this pre-Lehman elevation in big-bank credit spreads can most easily be viewed as an increase in the insolvency probabilities of big banks, while holding bail-out expectations at their high pre-Lehman levels.

For example, at a risk-neutral bail-out probability π of 0.6 and a loss in the event of insolvency and no bail-out of $L = 0.60$, the risk-neutral expected loss given default is $(1 - \pi)L = 0.24$. An increase in credit spreads of 60 basis points can then be interpreted as an increase in the risk-neutral annualized default probability of big banks in August 2007 of about 250 basis points. Berndt, Douglas,

Duffie, and Ferguson (2018) estimate that annualized risk-neutral default probabilities in mid-2000 were roughly double actual default probabilities, for firms with five-year CDS rates near those of the largest U.S. banks at that time. At this ratio, a 2.5% increase in risk-neutral default probability translates into an increase in the annualized actual default probability of big banks of roughly 1.25%, which seems a plausible investor reaction to the first hints of a major mortgage crisis. Big-bank default risk is highly systematic, so 2-to-1 probably understates the ratio of risk-neutral default probabilities to actual probabilities for big banks. This numerical example is not intended to be quantitatively precise, but rather to shed light on whether the August 2007 run-up in credit spreads could reasonably be related to changes in insolvency risk rather than changes in bailout expectations. We are not aware of any discussion in 2007 of changes in the perceived implicit support by the government of big banks. Major rating agencies maintained large sovereign uplifts in their credit ratings of big banks until after the failure of Lehman.

Table 3 shows the fitted pairs of pre-Lehman and post-Lehman bailout probabilities. Table 4 reports summary statistics across G-SIBs, D-SIBs and other firms of the decomposition of the market values of total future bank cash flows into various components, as fractions of the market value of equity. As fitted, solvency ratios improved significantly for big banks in the post-Lehman period.

Table 3: **Data-consistent pairs of bailout probabilities** This table reports our estimates for π_{pre}^G and π_{pre}^D , at various values for $\pi_{\text{post}}^G = \pi_{\text{post}}^D$.

π_{post}	π_{pre}^G	π_{pre}^D
0.30	0.67	0.51
0.20	0.63	0.46
0.10	0.60	0.43
0.00	0.54	0.39

Table A.2 in the appendix shows, for several fitted pairs of pre-Lehman and post-Lehman bailout probabilities, the average across G-SIBs of the market values of bank assets, total future bank cash flows and stakeholder claims, as well as the insolvency threshold, as fractions of the market value of equity. As fitted, with an increase in bailout probability, the estimated insolvency thresholds are lower. Appendix Table A.3 reports similar results for D-SIBs.

Depending on the assumed post-Lehman bailout probability, the market value y_3 of all future

Table 4: **Parameter estimates and firm value components** This table reports summary statistics for the calibrated model parameters k , $k + \phi$ and σ . It also shows the average and median solvency ratio d^s , LGL proxy l^1 and distance to default d . We also compute the risk-adjusted distance of the current market value of bank cash flows from that at the insolvency threshold, $d^y = [\log(Y(V_t)) - \log(Y(V^*))]/\sigma$. In the calibration, the bailout recapitalization achieves that debt is priced at par, that is, $B = P$. Post-Lehman bailout probabilities for big banks are set to 0.2, and pre-Lehman probabilities are as reported in Table 3.

	d^s	l^1	d	d^y	k	$k + \phi$	σ	d^s	l^1	d	d^y	k	$k + \phi$	σ
	Mean							Median						
G-SIB, pre	5.79	-2.69	3.10	0.42	3.25	0.56	0.21	5.77	-2.75	3.08	0.51	3.13	0.53	0.21
G-SIB, post	7.46	-5.02	2.44	1.20	0.98	0.24	0.14	6.83	-4.58	2.36	1.18	0.94	0.19	0.12
D-SIB, pre	5.52	-1.77	3.75	0.34	3.67	0.40	0.34	5.78	-1.90	3.91	0.36	3.72	0.30	0.33
D-SIB, post	7.99	-4.47	3.51	0.56	1.44	0.29	0.17	7.94	-3.98	3.72	0.63	1.47	0.21	0.12
Others, pre	4.78	-0.97	3.81	5.26	3.01	0.76	0.50	4.89	-0.90	3.91	5.28	2.95	0.60	0.48
Others, post	4.27	-0.74	3.53	5.24	1.13	0.53	0.39	4.30	-0.68	3.57	5.25	1.11	0.46	0.36

government subsidies ranges between 430% and 670% of the value of market equity. The market value of the next bailout subsidy alone, $U_r(x)\pi(\widehat{V} - V^*)$, is about one-half to two-thirds of the market value of all future subsidies, leaving the remaining y_3 as the market value of subsidies at subsequent bailouts. These subsidies are benefits to both creditors and equity shareholders. However, at the issuance of the debt, the creditors pay for the value of these subsidies in the form of higher bond prices (lower yield spreads), to the benefit of shareholders.

In Appendix D, we show that our findings are robust to an alternative measure of interest expense, and to alternative specifications of the price of debt at bailout and of the default boundary V^* .

8. Discussion and Concluding Remarks

As we have shown, high pre-Lehman bailout probabilities for big banks imply large subsidies of their debt financing costs. In order to maximize the market value of equity, these large banks would therefore have grown large balance sheets. Indeed, Figure 4 shows the rapid pre-Lehman growth of total assets of nine systemically important U.S. financial institutions, both banks and investment banks. After the crisis, we find that the reduced bailout prospects of big banks raised their cost of debt financing by roughly one half, at a given level of solvency. This in turn reduced the incentives of these firms to grow large balance sheets. As shown in Figure 4, post-Lehman asset growth is much more muted. Asset growth was also discouraged after the crisis by large increases in regulatory capital requirements, because, from the viewpoint of legacy equity shareholders, equity is a more costly source of financing

than debt.

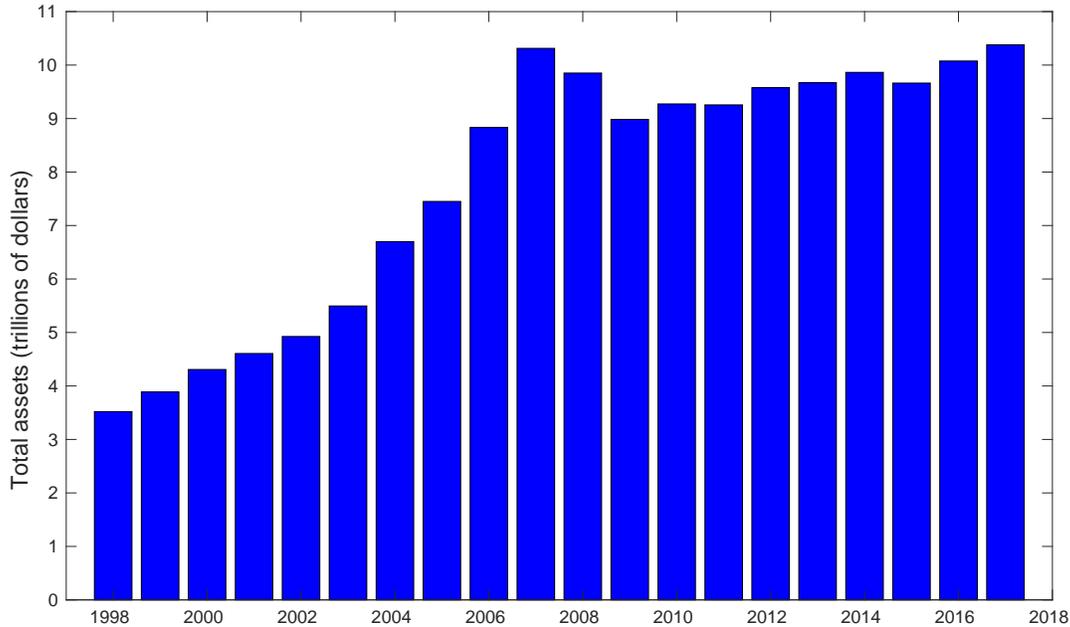


Figure 4: **Total assets of nine systemically important financial firms** The figure shows the total assets of the nine largest banks and investment banks—J.P. Morgan, Bank of America, Citibank, Wells Fargo, Goldman Sachs, Lehman Brothers, Merrill Lynch, and Bear Stearns. The data are sourced from the firms’ 10K filings.

Sarin and Summers (2016) suggest that the high credit spreads of large U.S. banks that prevailed in 2015 reflect high levels of default risk at that time, and that these firms were then about as likely to default as they were before the crisis. Our analysis suggests, instead, that high big-bank credit spreads in 2015, relative to pre-crisis years, are due to reduced bailout expectations. Buttressing our interpretation of the data, Figure 5 shows significant improvements by 2015 in the asset-weighted average solvency ratios of the largest U.S. financial institutions, the same firms whose assets are depicted in Figure 4. We define the solvency ratio of a firm to be the firm’s accounting tangible common equity divided by the estimated standard deviation of the annual change of its asset value. The improved solvency of large banks is due to significantly higher regulatory bank capital requirements, especially for G-SIBs. Rosengren (2014), Carney (2014), and Tucker (2014) describe the very large increases in capital buffers of the largest banks that were induced by post-crisis reform of bank capital regulations. In summary, the credit spreads of big banks were much higher after the crisis than before *despite* major increases in their capital buffers, and were apparently due to increased expected losses to creditors in the event of insolvency, rather than by high probabilities of insolvency. Our primary hypothesis that

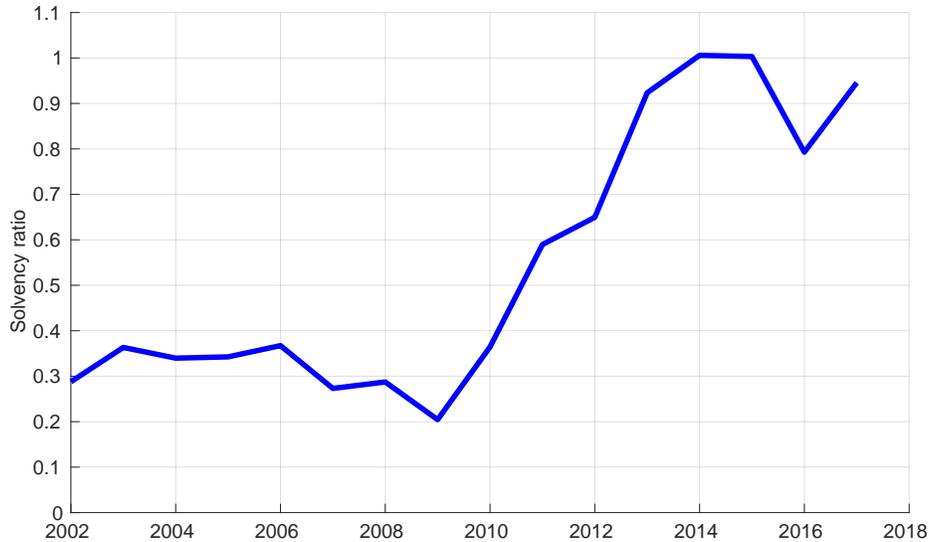


Figure 5: **Solvency ratios of the largest U.S. banks.** Tangible equity divided by an estimate of the standard deviation of the annual change in asset value. Asset-weighted average of J.P. Morgan, Bank of America, Citibank, Wells Fargo, Goldman Sachs, Lehman Brothers, Merrill Lynch, and Bear Stearns. Tangible equity data are obtained from 10K filings.

the post-crisis increase in expected insolvency losses are the result of a decline of “too-big-to-fail,” that is, lower reliance by big-bank creditors on the prospect of a government bailout.

An alternative, behavioral, explanation for high post-crisis big-bank credit spreads is that, before the crisis, big-bank creditors had little awareness that these banks could actually fail. The idea that market participants and regulators placed unduly low weight on the likelihood of a financial crisis is supported by [Gennaioli and Shleifer \(2018\)](#). By this line of reasoning, once Lehman failed and several other big banks had close calls, creditors could have become much more aware of risks of failure that were already elevated well before the crisis, but had been badly under-estimated. This would have caused wholesale bank credit spreads to remain elevated after the crisis. This story does not rely on changes in the likelihood of bailout, but rather on changes in the perceived likelihood of insolvency. The fact that post-crisis solvency buffers eventually got much higher than their pre-Lehman levels implies that the updating of insolvency probabilities after the crisis cannot reasonably have been based on a rational updating of beliefs based on new information.

If this alternative story applies, then the fact that big-bank credit spreads have remained high relative to their pre-Lehman levels would imply that the crisis-induced increase in the perception of

bank failure risk would need to have persisted for some years after the crisis. Historically, however, we are not aware of previous financial crises in which a large crisis-induced jump in wholesale big-bank credit spreads persisted well beyond the end of the crisis. For example, [Gorton and Tallman \(2018\)](#) use the “currency premium” as a gauge of wholesale bank (Clearinghouse) paper, around 19th century banking crises. Regarding the banking panic of 1893, for example, they write: “As gold inflows helped to restore reserve levels following suspension of convertibility on August 3, reports of redeposit of funds in New York Clearing House banks (presumably by interior correspondents) all contributed to an improvement to the financial setting. The key indicators for the banking system—the reserve deficit and the currency premium—become noticeably benign in newspaper articles. By August 31, the currency premium was less than one percent (0.625% in New York Tribune page 3, column 1). We find evidence from both the stock and bond markets that is consistent with the hypothesis.” This quote and the data analysis supporting it, shown in Figures 7 and 8 of [Gorton and Tallman \(2018\)](#), suggest that credit spreads jumped up during the 1893 crisis and then quickly went back down again within weeks after the panic.

In the setting of our research, were it not for a post-Lehman drop in the creditor-perceived probability π of a government bailout, we would have expected big-bank wholesale credit spreads to go back down (at a given level of solvency) with a general improvement in the economy and bank solvency. That is not what we find.

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A. Additional Figures and Tables

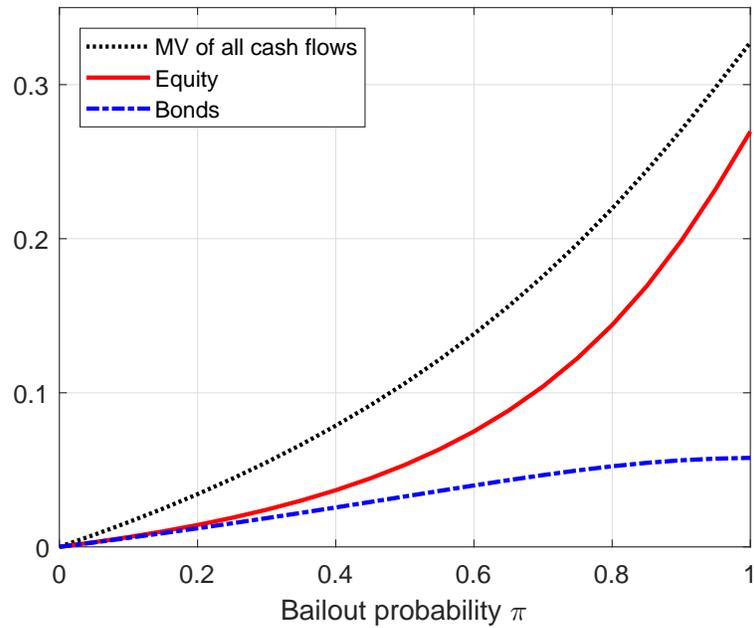


Figure A.1: **Effect of bailout on cash flow components.** The figure shows the increase in the market value (MV) of all cash flows, equity and bonds as the bailout probability π moves away from zero. There are no deposits. The other underlying primitive model parameters are $V_0 = 1$, $\sigma = 0.1317$, $\alpha = 0.5$, $r = 0.0230$, $k = 0.0221$, $\phi = -0.0136$, $\kappa = 0.35$, $m = 0.4233$, $c = 0.0680$, and $B = P = 0.5137$. These parameter choices are consistent with the calibrated model for Goldman Sachs in mid-September 2008, scaled to $V_0 = 1$.

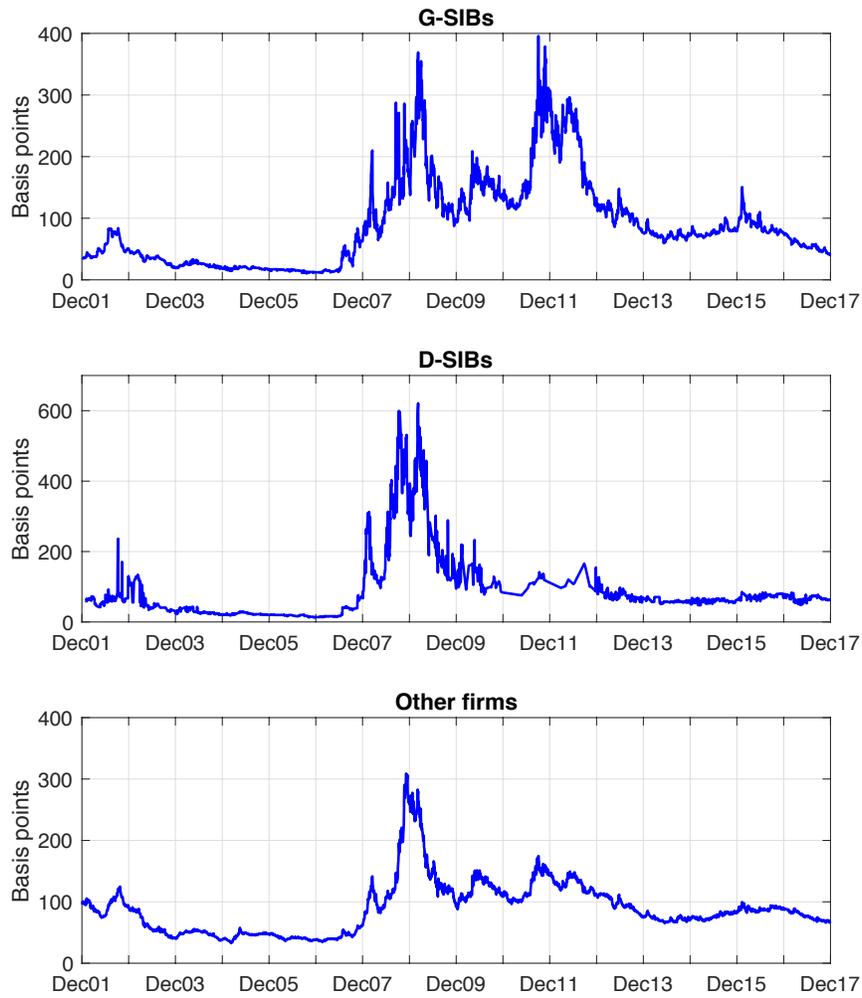


Figure A.2: **Median five-year CDS rates** The figure shows the daily times series of median five-year CDS rates. For G-SIBs and D-SIBs, only those days on which CDS rates are available for four or more firms are shown. For other firms, only days on which CDS rates are available for 50 or more firms are shown.

Table A.1: **Panel regression results for estimated bailout probabilities** This table reports the results for the panel data regression (7), when $\pi_{pre}^G = 0.63$ and $\pi_{post}^G = 0.2$ for G-SIBs; $\pi_{pre}^D = 0.46$ and $\pi_{post}^D = 0.2$ for D-SIBs; and $\pi_t^i = 0$ for non-bank firms i and dates t . CDS rates are measured in basis points. The distance to default and LGL proxies are computed as in (4) and (5). The benchmark sector is Basic Materials and the benchmark month is December 2017. Driscoll-Kraay standard errors that are robust to heteroskedasticity, autocorrelation and cross-sectional dependence are reported in parentheses. The data include 637 firms, over 2002–2017. The pre-Lehman period is from January 1, 2002 to September 15, 2008 and the post-Lehman period is from September 16, 2008 to December 31, 2017.

	Estimate	Std err
Constant	-1.539	0.023
Distance to default	-0.761	0.005
LGL proxy l^1	0.457	0.003
LGL proxy l^2	-0.140	0.001
Consumer Goods	0.044	0.005
Consumer Services	0.096	0.005
Energy	0.094	0.005
Financials	0.238	0.007
Healthcare	-0.070	0.006
Industrials	-0.123	0.005
Technology	0.139	0.010
Telecommunications Services	-0.013	0.008
Utilities	-0.025	0.003
Month FE	✓	
GSIBs and DSIBs FE	✓	
R^2	0.781	
RMSE	0.464	

Table A.2: Market value of bank cash flows and stakeholder claims for G-SIBs This table reports the market values of assets and of total net cash flows, and the market value of stakeholder claims, for the average G-SIB during the pre-Lehman period. The components are reported as fractions of market equity. V_0 is the current level of assets in place. V^* is the default threshold. $Y(V_0)$ is the total market value of all net cash flows when the current market value of assets in place is V_0 . y_3 is the market value of all future cash flows injected by the government, y_3^n is the market value of future bailout subsidies associated with the next bailout, and y_3^f is the market value of future bailout subsidies associated with the subsequent bailouts. v_1 is the total value of the claims of all current depositors, v_2 is the market value of all claims by current bondholders, and v_3 is the government's claim in return for all of its future successive bailout injections. D and P are the observed notional of deposits and bonds. The bailout recapitalization achieves that debt is priced at par, that is, $B = P$.

π_{post}	π_{pre}	V_0	V^*	$Y(V_0)$	y_3	y_3^n	y_3^f	v_1	v_2	v_3	D	P
<i>Pre-Lehman bailout probability is calibrated to data</i>												
0.30	0.67	20.43	9.90	6.50	6.68	3.94	2.74	1.75	3.08	0.67	2.56	3.08
0.20	0.63	20.25	10.03	6.43	5.67	3.52	2.15	1.77	3.08	0.59	2.56	3.08
0.10	0.60	20.15	10.09	6.40	5.15	3.31	1.84	1.78	3.08	0.54	2.56	3.08
0.00	0.54	20.03	10.20	6.34	4.34	2.92	1.42	1.80	3.08	0.47	2.56	3.08
<i>Bailout probabilities are set to zero</i>												
0.00	0.00	19.64	10.73	6.05	0.00	0.00	0.00	1.97	3.08	0.00	2.56	3.08

Table A.3: Market value of bank cash flows and stakeholder claims for D-SIBs This table reports the market values of assets and of total net cash flows, and the market value of stakeholder claims, for the average D-SIB during the pre-Lehman period. The components are reported as fractions of market equity. V_0 is the current level of assets in place. $Y(V_0)$ is the total market value of all net cash flows when the current market value of assets in place is V_0 . y_3 is the market value of all future cash flows injected by the government, y_3^n is the market value of future bailout subsidies associated with the next bailout, and y_3^f is the market value of future bailout subsidies associated with the subsequent bailouts. v_1 is the total value of the claims of all current depositors, v_2 is the market value of all claims by current bondholders, and v_3 is the government's claim in return for all of its future successive bailout injections. D and P are the observed notional of deposits and bonds. The bailout recapitalization achieves that debt is priced at par, that is, $B = P$.

π_{post}	π_{pre}	V_0	V^*	$Y(V_0)$	y_3	y_3^n	y_3^f	v_1	v_2	v_3	D	P
<i>Pre-Lehman bailout probability is calibrated to data</i>												
0.30	0.51	15.99	3.98	3.31	4.87	3.40	1.47	0.65	1.22	0.43	1.02	1.22
0.20	0.46	16.07	4.12	3.28	4.05	2.97	1.08	0.68	1.22	0.36	1.02	1.22
0.10	0.43	16.12	4.19	3.26	3.63	2.75	0.88	0.70	1.22	0.32	1.02	1.22
0.00	0.39	16.16	4.27	3.24	3.20	2.52	0.68	0.71	1.22	0.28	1.02	1.22
<i>Bailout probabilities are set to zero</i>												
0.00	0.00	16.06	4.82	3.13	0.00	0.00	0.00	0.90	1.22	0.00	1.02	1.22

B. Model Solution

This appendix provides explicit solutions for the basic model of valuation of bank debt and equity claims, including the asset default boundary V^* .

The conditional Laplace transform of the default time $\tau = \inf\{t : V_t \leq V^*\}$ evaluated at any positive constant θ is

$$U_\theta(V_t) = E^{\mathbb{Q}} \left(e^{-\theta(\tau-t)} | V_t \right) = \left(\frac{V_t}{V^*} \right)^{-\Gamma(\theta)},$$

where

$$\Gamma(\theta) = \frac{r - k - \sigma^2/2 + \sqrt{(r - k - \sigma^2/2)^2 + 2\theta\sigma^2}}{\sigma^2}.$$

In particular, the market value at any time $t < \tau$ of receiving one unit of account at the default time τ is $U_r(V_0)$.

B.1 Solution for the values of bank contingent claims, given the default boundary

When the current level of assets in place is x , we obtain the following market values of various respective contingent claims, taking V^* as given. First, the market value of the claim to all cash flows associated with a zero-debt version of the bank is

$$y_0(x) = x.$$

The market value of all future distress costs, including those associated with potential subsequent defaults, is

$$y_1(x) = U_r(x) \left[(1 - \pi)(1 - \alpha)V^* + \pi y_1(\widehat{V}) \right],$$

where, solving for the special case of $x = \widehat{V}$, we have

$$y_1(\widehat{V}) = \frac{U_r(\widehat{V})(1 - \pi)(1 - \alpha)V^*}{1 - \pi U_r(\widehat{V})}.$$

The market value of all future tax shields, making the most basic tax assumptions of [Leland \(1994a\)](#), is

$$y_2(x) = \kappa \frac{cP + dD}{r} (1 - U_r(x)) + \pi U_r(x) y_2(\widehat{V}),$$

where

$$y_2(\widehat{V}) = \frac{\kappa \frac{cP + dD}{r} (1 - U_r(\widehat{V}))}{1 - \pi U_r(\widehat{V})}.$$

The market value of all future bailout cash flows injected by the government (gross of the government's equity claims) is

$$y_3(x) = U_r(x) \pi \left[\widehat{V} - V^* + y_3(\widehat{V}) \right],$$

where

$$y_3(\widehat{V}) = \frac{\pi U_r(\widehat{V}) (\widehat{V} - V^*)}{1 - \pi U_r(\widehat{V})}.$$

The liquidation deposit guarantee requires cash flows from the government with a current market value of

$$y_4(x) = U_r(x) \left[(1 - \pi)(D - \alpha V^*)^+ + \pi y_4(\widehat{V}) \right],$$

where

$$y_4(\widehat{V}) = \frac{(1 - \pi) U_r(\widehat{V}) (D - \alpha V^*)^+}{1 - \pi U_r(\widehat{V})}.$$

Banks may also earn extra rents, or incur extra costs, on the asset side. The current market value of these extra rents or costs is

$$y_5(x) = q(x) - U_r(x) q(V^*) + \pi U_r(x) y_5(\widehat{V}),$$

where

$$q(x) = E \left[\int_0^\infty e^{-rt} \phi V_t dt \right] = \frac{\phi}{k} x$$

for some $\phi > -k$, and

$$y_5(\widehat{V}) = \frac{q(\widehat{V}) - U_r(\widehat{V})q(V^*)}{1 - \pi U_r(\widehat{V})}.$$

It is straightforward to show that

$$y_5(x) = q(x) - \left(\frac{q(V^*)}{1 - \pi U_r(\widehat{V})} - \frac{\pi q(\widehat{V})}{1 - \pi U_r(\widehat{V})} \right) U_r(x).$$

The total market value of all cash flows available to the firm's current claimants is thus

$$Y(x) = y_0(x) - y_1(x) + y_2(x) + y_3(x) + y_4(x) + y_5(x). \quad (\text{B.1})$$

The market value of the claims of current depositors is

$$v_1(x) = D \frac{d}{r} (1 - U_r(x)) + U_r(x) \left[\pi v_1(\widehat{V}) + (1 - \pi) D \right],$$

where

$$v_1(\widehat{V}) = \frac{D \frac{d}{r} (1 - U_r(\widehat{V})) + (1 - \pi) D U_r(\widehat{V})}{1 - \pi U_r(\widehat{V})}.$$

Extending the integration-by-parts argument of [Leland \(1994a\)](#), the market value of the claims of current bondholders is,

$$v_2(x) = W(1 - U_{r+m}(x)) + U_{r+m}(x) (\pi B + (1 - \pi)(\alpha V^* - D)^+), \quad (\text{B.2})$$

where

$$W = P \frac{c + m}{r + m}$$

is the total market value of bonds that are default free but otherwise equivalent to those issued by the bank. We always take $B \leq W$ because it is impossible for the bonds to be worth more than their risk-free value W . In the calibration, we also use the liquidation bond recovery $\min((\alpha V^* - D)^+, B)$ in place of $(\alpha V^* - D)^+$, but for notational simplicity stick to the latter in this section.

The condition $v_2(\widehat{V}) = B$ implies that

$$\widehat{V} = h(V^*) \equiv V^* \left(\frac{W - B}{W - \pi B - (1 - \pi)(\alpha V^* - D)^+} \right)^{-\frac{1}{\eta}}, \quad (\text{B.3})$$

where $\eta = \Gamma(r + m)$.

In return for all of its future successive bailout injections, the government has a claim with a market value of

$$v_3(x) = U_r(x) \pi \left[H(\widehat{V}) + v_3(\widehat{V}) \right],$$

where

$$v_3(\widehat{V}) = \frac{\pi U(\widehat{V}) H(\widehat{V})}{1 - \pi U(\widehat{V})}.$$

The total market value of all claims on the bank's net future cash flows is equal to the market value of total cash flows available, so the market value of the bank's equity is

$$H(x) = Y(x) - v_1(x) - v_2(x) - v_3(x), \quad x \geq V^*. \quad (\text{B.4})$$

By definition, $H(x) = 0$ for $x < V^*$. We can rewrite Equation (B.4) as

$$H(x) = \left(1 + \frac{\phi}{k} \right) x + a + b(V^*) U_{r+m}(x) + g(V^*) U_r(x), \quad x \geq V^*, \quad (\text{B.5})$$

where

$$\begin{aligned}
a &= \kappa \frac{cP + dD}{r} - \frac{dD}{r} - W \\
b(V^*) &= W - \pi B - (1 - \pi)(\alpha V^* - D)^+ \\
g(V^*) &= \pi B + (1 - \pi)(\alpha V^* - D)^+ - \kappa \frac{cP + dD}{r} + D \frac{d}{r} - \left(1 + \frac{\phi}{k}\right) V^*.
\end{aligned}$$

B.2 Optimal Default Boundary

For the case of an endogenous insolvency boundary level of assets V^* that maximizes shareholder value, V^* can be conjectured and then verified from the smooth pasting condition, namely that the market value of equity is continuously differentiable at V^* , implying that

$$H'(V^*) = 0. \quad (\text{B.6})$$

The smooth-pasting condition (B.6) reduces to

$$0 = \left(1 + \frac{\phi}{k}\right) V^* - \eta b(V^*) - \gamma g(V^*), \quad (\text{B.7})$$

where $\gamma = \Gamma(r)$. Expression (B.7) provides an equation for the default boundary V^* that we can now solve explicitly, thus providing explicit solutions for \widehat{V} and, in turn, all of the contingent claim market valuation functions $y_1, y_2, y_3, y_4, y_5, v_1, v_2, v_3$, and H that we have considered.

Rewriting Equation (B.7) yields

$$\begin{aligned}
(1 + \gamma) \left(1 + \frac{\phi}{k}\right) V^* &= \eta [W - \pi B - (1 - \pi)(\alpha V^* - D)^+] \\
&\quad + \gamma \left[(1 - \pi)(\alpha V^* - D)^+ - \kappa \frac{cP + dD}{r} + D \frac{d}{r} + \pi B \right].
\end{aligned} \quad (\text{B.8})$$

There are two cases to consider when solving for V^* :

Case 1: If $\alpha V^* - D \leq 0$, then Equation (B.8) can be re-written as

$$V^* = \frac{\eta(W - \pi B) + \gamma \left(-\kappa \frac{cP + dD}{r} + D \frac{d}{r} + \pi B\right)}{(1 + \gamma) \left(1 + \frac{\phi}{k}\right)}. \quad (\text{B.9})$$

We obtain a solution $V^* > 0$ in (B.9) as long as

$$0 < \frac{\eta(W - \pi B) + \gamma \left(-\kappa \frac{cP+dD}{r} + D \frac{d}{r} + \pi B \right)}{(1 + \gamma) \left(1 + \frac{\phi}{k} \right)} \leq \frac{D}{\alpha}. \quad (\text{B.10})$$

Case 2: If $\alpha V^* - D > 0$, Equation (B.8) yields

$$V^* = \frac{\eta(W - \pi B) + \gamma \left(-\kappa \frac{cP+dD}{r} + D \frac{d}{r} + \pi B \right) - (1 - \pi)(\gamma - \eta)D}{(1 + \gamma) \left(1 + \frac{\phi}{k} \right) - \alpha(1 - \pi)(\gamma - \eta)}. \quad (\text{B.11})$$

We obtain a solution $V^* > 0$ in (B.11) as long as

$$\frac{\eta(W - \pi B) + \gamma \left(-\kappa \frac{cP+dD}{r} + D \frac{d}{r} + \pi B \right) - (1 - \pi)(\gamma - \eta)D}{(1 + \gamma) \left(1 + \frac{\phi}{k} \right) - \alpha(1 - \pi)(\gamma - \eta)} > \frac{D}{\alpha}. \quad (\text{B.12})$$

Note that $(1 + \gamma) \left(1 + \frac{\phi}{k} \right) > 0$ and $(1 + \gamma) \left(1 + \frac{\phi}{k} \right) - \alpha(1 - \pi)(\gamma - \eta) > 0$. The conditions (B.10) and (B.12) are mutually exclusive. As a results, there exists a unique solution for V^* if and only if $\eta(W - \pi B) + \gamma \left(-\kappa \frac{cP+dD}{r} + D \frac{d}{r} + \pi B \right) > 0$.

B.3 Extending the model to incorporate liability insurance assessments

Our model can accommodate deposit and other liability insurance premiums, as follows. We suppose a deposit insurance rate i_D and an insurance assessment rate i_P on other liabilities. Given V^* , the total market value of future insurance premia is

$$y_6(x) = \frac{P i_P + D i_D}{r} (1 - U_r(x)) + U_r(x) \pi y_6(\widehat{V}).$$

As with the other value components, this implies a linear equation for $y_6(\widehat{V})$, and thus an explicit solution for $y_6(x)$ at any x . The valuation of total available bank cash flows is now extended from (B.1) to

$$Y(x) = y_0(x) - y_1(x) + y_2(x) + y_3(x) + y_4(x) + y_5(x) - y_6(x).$$

From this point, the solution method for the boundary V^* and, from that, all of the value components, is just as for the basic model. The solution is again explicit.

C. A Model With Senior Bonds and Bail-in Junior Bonds

This appendix generalizes the basic model to allow for senior and junior bonds, with the objective of separate identification of the risk-neutral probabilities π of bail-out, ψ of bail-in, and $1 - \pi - \psi$ of liquidation at bankruptcy. The dynamic Equation (3) for V_t is maintained. We have the same non-bond parameters $D, d, r, \sigma, k, \alpha$ and κ as for the basic version of the model. The senior and junior bonds have the same maturity parameter m . As with the basic model, the senior bonds have principal P and coupon rate c . At a bailout, the assets in place are increased to some level \widehat{V} by a capital injection that increases the market value of the senior bonds to B , as before. The junior bonds have principal J , coupon rate j , and at a bailout have whatever market value is implied by the capital injection. At a bail-in, the junior bonds are given all of the equity in the bank, and the bank emerges with only its original senior bonds. From that point, for simplicity, we assume that only bailout and liquidation are possible. The values of all elements of the capital structure are then given by the basic version of the model. The default boundary V^* for the basic model without junior debt will therefore apply after a bail-in. This is different from the default boundary V^* that initially applies when there is bail-in junior debt. An alternative and more complicated version of the model would have a bail-in design that restructures the liabilities so as to introduce after bail-in a given new amount of senior and junior bonds.

When any existing bond matures at time t , the same principle amount of the same type of debt is issued at its current market value, which could be at a premium or discount to par depending on V_t . Newly issued senior and junior bonds have the original coupon rates c and j , respectively. The original exponential maturity distribution is always maintained.

The primitive parameters of the model are $c, P, B, J, j, m, D, d, r, \sigma, k, \alpha$ and κ .

We first take the default boundary V^* and first bail-out level \widehat{V} for assets in place as given, and later derive the associated value-consistency and smooth-fit condition determining these two boundaries.

When the current level of assets in place is x , we obtain the following market values of various respective contingent claims. First, the market value of the claim to all cash flows associated with a zero-debt version of the bank is

$$y_0(x) = x.$$

The market value of all future distress costs, including those associated with potential subsequent defaults, is

$$y_{1b}(x) = U_r(x)[(1 - \pi - \psi)(1 - \alpha)V^* + \pi y_{1b}(\widehat{V}) + \psi y_1(V^*)],$$

where $y_1(\cdot)$ is the solution for the value of distress costs in the basic model for a bank with parameters $c, P, B, m, D, d, r, \sigma, k, \alpha, \pi$ and κ . This equation implies an explicit solution for $y_{1b}(\widehat{V})$.

The market value of all future tax shields is

$$y_{2b}(x) = \kappa \frac{Pc + Dd + Jj}{r} (1 - U_r(x)) + U_r(x)[\pi y_2(\widehat{V}) + \psi y_2(V^*)].$$

This implies an explicit solution for $y_{2b}(\widehat{V})$.

The market value of all future cash flows injected by the government, before considering the effect of government equity claims, is

$$y_{3b}(x) = U_r(x)[\pi (\widehat{V} - V^* + y_{3b}(\widehat{V})) + \psi y_3(V^*)].$$

Again, we have an explicit solution for $y_{3b}(\widehat{V})$.

The liquidation deposit guarantee requires cash flows from the government with a current market value of

$$y_{4b}(x) = U_r(x) \left[(1 - \pi - \psi) (D - \alpha V^*)^+ + \pi y_{4b}(\widehat{V}) + \psi y_4(V^*) \right].$$

Again, we have an explicit solution for $y_{4b}(\widehat{V})$.

The total market value of all net cash flows available to the firm's *current* claimants is

$$Y_b(x) = y_0(x) - y_{1b}(x) + y_{2b}(x) + y_{3b}(x) + y_{4b}(x).$$

For simplicity, we ignore extra rents or costs on the asset side as well as liability insurance premia.

The total value of the claims of all current depositors is

$$v_{1b}(x) = D \frac{d}{r} (1 - U_r(x)) + U_r(x) \left(\pi v_{1b}(\widehat{V}) + (1 - \pi - \psi)D + \psi v_1(V^*) \right).$$

We can solve explicitly for $v_{1b}(\widehat{V})$.

The market value of all claims by current senior bondholders is

$$\begin{aligned} v_{2b}(x) = & \frac{cP + mP}{r + m} (1 - U_{r+m}(x)) \\ & + U_{r+m}(x) [\pi B + (1 - \pi - \psi) \max(\alpha V^* - D)^+, P] + \psi v_2(V^*). \end{aligned}$$

We have the consistency condition

$$v_{2b}(\widehat{V}) = B, \tag{C.1}$$

which determines \widehat{V} uniquely given V^* .

In return for all of its future successive bailout injections, the government has a claim with a market value of

$$v_{3b}(x) = U_r(x) [\pi G(\widehat{V}) + \pi v_{3b}(\widehat{V}) + \psi v_3(V^*)],$$

where $G(x)$ is the equity value of the original bank with assets in place of x .

The market value of all claims by current junior bondholders is

$$v_{4b}(x) = \frac{Jj + mJ}{r + m} (1 - U_{r+m}(x)) + U_{r+m}(x) \left[\pi v_{4b}(\widehat{V}) + (1 - \pi - \psi)(\alpha V^* - D - P)^+ + \psi H(V^*) \right].$$

By definition, $G(x) = 0$ for $x \leq V^*$. The total market value of all claims on the bank's net future cash flows is equal to the market value of total cash flows available, so

$$G(x) = Y_b(x) - v_{1b}(x) - v_{2b}(x) - v_{3b}(x) - v_{4b}(x), \quad x \geq V^*. \tag{C.2}$$

Given the assets in place \widehat{V} after the first bailout, the default boundary V^* that maximizes share-

holder value can be conjectured and then verified from the smooth-pasting condition, namely that the market value of equity is continuously differentiable at V^* , implying that

$$\mathcal{G}(V^*, \widehat{V}) \equiv G'(V^*) = 0. \quad (\text{C.3})$$

We can calculate $\mathcal{G}(V^*, \widehat{V})$ explicitly as a function of V^* and \widehat{V} . We have reduced the solution of equilibrium for the model to the two equations (C.1) and (C.3) to solve for the two boundaries V^* and \widehat{V} .

D. Robustness Checks

In this appendix, we report results for three alternative implementations of the model calibration. First, we consider an alternative definition of the interest expense, IE , and use the Compustat quarterly total interest expenses. Second, we consider a specification where B_{it} is chosen to achieve a given post-bailout bond yield spread s , by setting $B_{it} = P_{it}(c_{it} + m_{it})/(r_{it} + s + m_{it})$. This is based on the idea that a government bailout would target a given level of creditworthiness of a large bank as judged in wholesale credit markets. We set $s = 100$ basis points. Third, we also consider the case when the liquidation bond recovery rates are exogenously specified as 40%.

Table D.1: Bailout probability estimates based on alternative calibration assumptions This table reports our estimates for π_{pre}^G and π_{pre}^D , at various values for $\pi_{\text{post}}^G = \pi_{\text{post}}^D$. The left panel presents the estimates when COMPUSTAT interest expense data are used to calibrating the cash flow rate $k + \phi$. The middle panels shows the estimates when B_{it} is specified as $B = P(c + m)/(r + s + m)$, with s equal to 100 basis points. The right panel presents the results for the case where loss given liquidation is set exogenously to 60%.

Accounting-based IE			Alternative specification of B			Exogenous bond recovery rates		
π_{post}	π_{pre}^G	π_{pre}^D	π_{post}	π_{pre}^G	π_{pre}^D	π_{post}	π_{pre}^G	π_{pre}^D
0.30	0.64	0.52	0.30	0.62	0.50	0.30	0.65	0.51
0.20	0.61	0.49	0.20	0.59	0.43	0.20	0.61	0.45
0.10	0.58	0.45	0.10	0.54	0.40	0.10	0.56	0.42
0.00	0.52	0.40	0.00	0.49	0.36	0.00	0.51	0.39

E. Model Calibration

The key input variables for the model calibration and their sources are described in Table E.1. We also list the model outputs and the restrictions through which they are identified.

Table E.1: **Model input and output variables** The table describes the input and output variables for the model calibration and explains how they are sourced or calibrated.

Parameter	Definition	Source/Value/Identifying assumption
Inputs		
H	Market value of equity	CRSP
P	Sum of long- and short-term debt	Compustat and 10-K/Q
D	Deposits	Compustat and 10-K/Q
$1/m$	Notional-weighted maturity	Compustat and 10-K
r	Risk-free rate	Constant-maturity Treasury rate, linearly interpolated at average bond maturity
r^d	Deposit rate	Computed from Compustat and 10-K/Q
c	Par coupon rate	Risk-free rate plus CDS rate minus CDS-bond basis
$1 - \alpha$	Fractional bankruptcy cost at liquidation	0.50
κ	Corporate tax rate	0.35
CD	Cash dividend	Compustat
IE	Interest expense for long- and short-term debt and deposits	Computed as $Dr^d + P(r + \text{CDS} - \text{basis})$
Outputs		
V^*	Assets boundary	Smooth-pasting condition
V	Assets in place	Model-implied market value of debt matches observed market value of debt
σ	Asset volatility	$\sigma_p = \frac{1}{\sqrt{h}} \text{StdDev}(\log(V_{t+h}) - \log(V_t) t \in p)$
ϕ	Extra asset rents/costs	Model-implied market equity matches observed market equity
$k + \phi$	Asset payout rate	For $t \in p$, ensure that $\text{Mean}\{k_t + \phi_t t \in p\}$ is equal to $\text{Mean}\{(CD_t + IE_t)/V_t t \in p\}$

For a given bank i , period p and bailout probabilities $\pi_{it} = \pi_{ip}$ for $t \in p$, the calibration proceeds as follows:

1. Make an initial guess of ρ_{ip} , which determines the asset drift parameter k_{it} via $k_{it} = \rho_{ip}r_{it}$, and asset volatility σ_{ip} using accounting data.²⁶ Set $\rho_{it} = \rho_{ip}$ and $\sigma_{it} = \sigma_{ip}$ for all $t \in p$.
2. For each date $t \in p$, we find the default threshold V_{it}^* as the solution to $H'(V_{it}^*) = 0$ by solving Equation (B.7).²⁷

²⁶The initial input parameters k_{ip} and σ_{ip} are computed from accounting data as

$$\rho_p = \frac{\text{Mean}\{(CD_t + IE_t)/BVA_t | t \in p\}}{\text{Mean}\{r_t | t \in p\}} \quad \text{and} \quad \sigma_p = \frac{1}{\sqrt{h}} \text{StdDev}\{\log(BVA_{t+h}) - \log(BVA_t) | t \in p\},$$

where BVA denotes book assets, h measures daily time steps, and the subscript i has been dropped from the notation.

²⁷For robustness checks where the liquidation bond recovery rate is exogenously given, we find V_{it}^* as the solution to a

3. For each date $t \in p$, find x_{it} such that $v_2(x_{it})$ computed according to Equation (B.2) matches the observed par value of debt.
4. For each date $t \in p$, find ϕ_{it} such that $H(x_{it})$ computed according to Equation (B.4) matches the observed market value of equity.
5. Re-compute ρ_{ip} and σ_{ip} as

$$\rho_p = \frac{\text{Mean}\{(CD_t + IE_t)/V_t | t \in p\} - \text{Mean}\{\phi_t | t \in p\}}{\text{Mean}\{r_t | t \in p\}},$$

$$\sigma_p = \frac{1}{\sqrt{h}} \text{StdDev}\{\log(x_{i,t+h}) - \log(x_{it}) | t \in p\},$$

after dropping the subscript i . If ρ_{ip} or σ_{ip} have changed, return to Step 2. Otherwise, stop.

fixed liquidation bond recovery rate $(\alpha V_{it}^* - D_{it})/P_{it}$.