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## Risk Taking and Interest Rates: Evidence from Decades in the Global Syndicated Loan Market

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# Risk Taking and Interest Rates: Evidence from Decades in the Global Syndicated Loan Market

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## Abstract

We study how low interest rates in the United States affect risk taking in the market of cross-border leveraged corporate loans. To the extent that actions of the Federal Reserve affect U.S. interest rates, our analysis provides evidence of a cross-border spillover effect of monetary policy. We find that before the crisis, lenders made ex-ante riskier loans to non-U.S. borrowers in response to a decline in short-term U.S. interest rates, and, after it, in response to a decline in longer-term U.S. interest rates. Economic uncertainty and risk appetite appear to play a limited role in explaining ex-ante credit risk. Our results highlight the potential policy challenges faced by central banks in affecting credit risk cycles in their own jurisdictions.

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# 1 Introduction

In the aftermath of the global financial crisis, low interest rates maintained by central banks in advanced economies have sparked an intense debate on cross-border spillover effects of monetary policy. One strand of the literature highlights international bank lending as a potential channel through which monetary conditions in one country affects capital flows and credit growth in the international financial system (Cetorelli and Goldberg (2011), Rey (2015), Bruno and Shin (2015), Obstfeld (2015)). While most studies in this literature focus on the direction and magnitude of cross-border financial flows, little has been done on studying the riskiness of these credit flows. Indeed, recent studies—for example, Kalemli-Ozcan, Sorensen, and Yesiltas (2012)—argues that quantitative measures of risky assets may not be sufficient in detecting financial vulnerability as risk taking in the financial sector could involve quality rather than the quantity of assets.[1]

Using market-based measures of credit risk and accounting for a broad set of determinants of risks of cross-border loans, this paper examines how changes in U.S. interest rates affect risk taking by financial institutions in the global syndicated term loan market. More specifically, we study whether low interest rates in the U.S. encourage lenders to supply riskier credit to non-U.S. borrowers, including borrowers in emerging market economies (EMEs), and whether this effect is more prominent for bank or non-bank lenders. In addition, we distinguish the differential responses of risk taking to changes in short-term and long-term U.S. interest rates. To the extent that actions of the Federal reserve affect short- and long-term U.S. interest rates, we provide evidence of a cross-border spillover effect of monetary policy - one that is associated with the global lending channel.

The theoretical literature presents a mixed view on the relationship between the real interest rate and risk taking in the financial sector, and most studies focus on depository institutions. In these models, the presence of asymmetric information and limited liability

[1] Kalemli-Ozcan, Sorensen, and Yesiltas (2012) present stylized facts on bank and firm leverage during the period 2000-09 using internationally comparable micro level data from many countries. Their results suggest that excessive risk taking before the crisis was not easily detectable because the risk involved the quality rather than the quantity of assets.

leads to a risk-shifting effect when interest rate rises, with an increase in deposit rate exacerbating agency problem and encouraging bank risk taking (see for example, Dell'Ariccia, Laeven, and Marquez (2014), Acharya and Viswanathan (2011)). In contrast, the portfolio allocation theory predicts that a fall in interest rate on safe assets induces investors to shift away from safe assets towards riskier ones, thereby raising the riskiness of their portfolio. Furthermore, for financial institutions with fixed rate obligations, such as life insurance companies and pension funds, the choice of risk taking may reflect a "search for yield" incentive which induces them to take on more risk as interest rates fall (Rajan (2005)). Thus, the overall effect of interest rates on risk taking depends on the nature of interest rate changes—short-term versus long-term—and the type of financial institutions in consideration—banks versus nonbanks.

This paper studies the empirical relationship between interest rates and risk taking along these dimensions. To capture risk taking in the broad financial sector, we look at risk characteristics of syndicated term loans originated in the global market over the last two decades during which co-movements between short-term and long-term interest rates exhibit different patterns before and after the global financial crisis. The global syndicated term loan market is a suitable venue to study this topic for several reasons. First, this market is comparable in size to the global corporate bond market and provides trillions of U.S. dollars a year to mostly nonfinancial corporations. Second, the vast majority of syndicated term loans are made to speculative-grade and obscure, non-rated corporations, so lender participants take on very large positions in risky leveraged assets. Third, a significant portion of the non-U.S. borrowers in the market is from EMEs, which allows us to identify the cross-border spillover effects. Fourth, while banks tend to originate the majority of syndicated term loans, they sell these loans very quickly to nonbank lenders, appearing to accommodate nonbanks' investment choices (Aramonte, Lee, and Stebunovs (2015)). Hence, the riskiness of these loans often reflects risk taking behavior by the broad financial system rather than just banks.

We focus on an ex-ante measure of credit risk proxied by loan spreads that are fixed at loan origination for the duration of a loan. As argued in Dell'Ariccia, Laeven, and Marquez (2014), ex-ante measures of risk taking have a couple of advantages. It helps alleviates concerns about endogenity of interest rate changes. It also allows us to focus on ex-ante risk perception of investors at loan origination rather than ex-post loan performance which could be affected by changes in credit quality of outstanding loans. To illustrate that loan spreads are a reliable proxy for ex-ante credit risk, for a subsample of loans we provide evidence that loan spreads are strongly correlated with probabilities of borrower default over a one-year horizon (a Basel II-consistent risk parameter) reported by lenders. Given that a typical loan has a maturity of several years, the loan spread captures the loan's credit risk smoothed (averaged) through the cycle.

Empirical studies on the link between interest rate and risk taking are often subject to two major identification challenges. First, interest rates are endogenous to economic conditions, so an estimated relationship between interest rates and ex-ante credit risk may reflect some latent factors that drive movements in the two variables simultaneously. Second, changes in credit risk of a loan may reflect responses to both demand and supply factors, and disentangling one from another is difficult.

We adopt the following strategies to address the identification concerns. First, our main focus is international spillover effects of risk taking, so we restrict our sample of loans to those originated to non-U.S. borrowers. This allows us to separate interest rate and credit cycles, and alleviate endogeneity concerns in relation to U.S. interest rates . As in Jimenez, Ongena, Peydro, and Saurina (2014) and Ioannidou, Ongena, and Peydro (2015), monetary conditions in one country are generally considered to be exogenous to developments in another country. Second, we include various variables that capture economic uncertainty and risk appetitive to control for influences from global factors. Third, we conduct our main analysis at the syndicated loan level, and introduce controls for both credit supply (push) factors and credit demand (pull) factors. More specifically, we use lender country- and lender type-year effects to control for credit supply cycles specific to each lender, and borrower-type year and borrower-region year fixed effects to control for credit demand cycles specific to each borrower industry and region in our sample.

We find robust evidence that lower interest rates in the U.S. encourage lenders to extend riskier loans in the global syndicated loan market, although the risk taking behavior responds differently to interest rate changes before and after the crisis. In the pre-crisis period, all else held equal, a decline in U.S. short-term interest rates encouraged lenders of various types and nationalities to originate riskier loans to non-U.S. borrowers. Once this effect is accounted for, movements in long-term interest rates did not have statistically significant effect on risk taking. Given that the Federal Reserve explicitly targeted the federal funds rate as a policy instrument during this period, our results can be seen as presenting evidence of a risk taking channel of monetary policy.[2]

In contrast, for the post-crisis period, our results suggest that lenders of various types and nationalities tended to originate riskier loans to non-U.S. borrowers in response to a decline in long-term interest rates. Notably, short-term rates were close to zero lower bound during this period, thus the post-crisis results are implicitly conditional on short-term rates being low. Since the implicit target of monetary policy was the longer-term rates (with the U.S. 10-year Treasury rate being a de facto target) during this period, our results again confirm the risk taking channel of monetary policy, although the operating channel is different from that in the pre-crisis period.[3] The response of risk taking to the long-end of the yield curve may reflect the return-on-safer-assets and search-for-yield motives discussed in the theoretical literature.

We acknowledge that one cannot contribute all the movement in interest rates to Federal Reserve's actions. For instance, following the European sovereign crisis, increased uncertainty and safe haven inflows to safer U.S. assets put additional downward pressures on U.S. longer-term interest rates. To control for these global factors, we include the European sovereign spread and economic uncertainty index in our regressions and find that these

[2]The Fed was very successful in achieving the target, with the spread between the effective federal rate and its target averaged just a couple of basis points with a low variance.

[3]Recent research by Krishnamurthy and Vissing-Jorgensen (2011), D'Amico, English, López-Salido, and Nelson (2012), and D'Amico and King (2013), suggests that the Federal Reserve's unconventional monetary policy actions have reduced U.S. longer-term interest rates on safer assets.

global factors played a limited role in explaining ex-ante credit risk of the syndicated loans. The effects of interest rates on risk taking remain robust.

Because investors often make risk management decisions based on credit risk of overall loan portfolios, not just individual loans, we also estimate risk taking responses to interest rates and other factors at a lender portfolio level. The results from these sets of regressions reinforce what we find from loan level regressions. In addition, we study how changes in interest rates affect risk taking by bank and nonbank financial lenders differently. We find that while banks took more risk in response to both a drop in short-term interest rates before the crisis and a decline in long-term rates after the crisis, nonbanks increased risk taking only when long-term rates decrease in the post-crisis period.

Our paper makes three important contributions to the literature on capital flows and the risk taking channel of monetary policy. First, we examine international spillover of monetary policy by studying the quality of cross-border credit flows rather than the quantity. Our findings present evidence of risk taking spillovers with low interest rates in one country encouraging lenders to originate riskier loans to borrowers in other countries. Such spillovers highlight the potential challenges faced by central banks, particularly those of EMEs, in affecting credit cycles in their respective jurisdictions. Second, we distinguish the different roles played by short-term and long-term interest rates in affecting the risk taking channel of monetary policy. Such a distinction allows us to compare the transmission channels of monetary policy before and after the crisis. Third, our analysis covers the broad financial system including both banks and nonbanks, which allows us to distinguish how risk taking by different financial financial institutions responds to changes in interest rates. Understanding the risk taking behavior by nonbank institutions is particularly important given the rapid expansion of this sector in the post-crisis period.

The reminder of the paper is organized as follows. Section 2 presents a brief literature review. We then describe the data source, outline key features of syndicated loans, and demonstrate that loan spreads are reasonable proxies for ex-ante credit risk metrics in section 3. Section 4 presents the empirical methodology used to examine the relationship between interest rates and ex-ante credit risk in the cross-border term loan market. Section 5 summarizes the empirical results and discusses a number of robustness tests. We conclude in section 6 with a few remarks on the implications of our findings for financial stability issues.

## 2 Literature review

Our paper contributes to several strands of the literature, including those on global lending channels and capital flows, as well as the risk taking channel of monetary policy.

Existing papers on capital flows and international bank lending mostly focus on the direction and magnitude of cross-border financial flows. Using aggregate banking statistics data, Bruno and Shin (2015) study the effects of monetary policy spillovers on cross-border capital flows. Their results suggest that an expansionary shock to U.S. monetary policy leads to a increase in cross-border bank capital flows and an increase in the leverage of international banks. Similarly, Rey (2015) investigates the determinants of the global financial cycle using aggregate capital flow data, and finds that a decline in the federal fund rate leads to an increase in EU bank leverage (measured as the ratio of assets and liabilities) and an expansion of global credit flows. Furthermore, Temesvary, Ongena, and Owen (2015) examine the global bank lending channel using U.S. bank-level data. Their analysis shows that in response to changes in U.S. monetary conditions—both conventional and unconventional—U.S. banks adjust their cross-border claims. In contrast, our paper studies cross-border monetary policy spillovers using a rich loan-level dataset, which enables us to properly control for different push and pull factors. More importantly, we focus on the quality instead of the quantity of of credit flows, and provide direct evidence of the global risk taking channel of monetary policy.

Our paper also complements the literature on the risk taking channel of monetary policy, which has been growing in recent years, in part, because of prevalence of low interest rates around the world. Most empirical studies focus on banks as credit risk takers and on short-term rates, typically associated with banks' funding costs, as the main driver. For instance, Dell'Ariccia, Laeven, and Suarez (2016) study how monetary policy affects risk taking in the U.S. banking system. Using confidential data on banks' internal ratings on loans to business, they find that ex-ante risk taking by banks is negatively associate with short-term interest rates and this negative relationship is more pronounced for highly capitalized banks. Because U.S. monetary policy is endogenous to U.S. credit and business cycles, they paper focuses on new business loans only which they argue are less likely to inform FOMC decisions.

A few papers explore the international dimension of risk taking of monetary policy and use it as an strategy to strengthen identification. The findings in Jimenez, Ongena, Peydro, and Saurina (2014) show that banks in Spain take more credit risk as euro-area short-term interest rates decline. In particular, such declines induce worse capitalized banks, on the extensive margin, to grant more loan applications to ex-ante risky firms and, on the intensive margin, to commit larger loan volumes with fewer collateral requirements to these firms. The study by Ioannidou, Ongena, and Peydro (2015) provide evidence that banks in Bolivia grant riskier loans as U.S. short-term interest rates decline. Specifically, a decrease in the federal funds rate prior to loan origination raises the monthly probability of default on individual bank loans but banks do not seem to price this additional risk adequately. In both papers, monetary policy is exogenous to the host country of banks in consideration (Spain is in the euro area, where the European Central Bank determines a monetary policy stance, and Bolivia is a small, dollarized economy, with negligible importance for U.S. monetary policy). We follow a similar identification method by examining the relationship between U.S. interest rates and the ex-ante credit risk of loans extended to non-U.S. borrowers. However, our data of cross-border loans covers lenders and borrowers from a wide range of countries, including EMEs, which allows us to look at risk taking in the global market. Moreover, we examine the risk taking by both bank and nonbank lenders, and differentiate the effects of movements in short-term and long-term interest rates.

Our paper is most closely related to Aramonte, Lee, and Stebunovs (2015), who consider longer-term interest as a main driver for ex-ante credit risk taking by a variety of financial intermediaries. They use U.S. syndicated loan data over the zero-lower bound period to show that, in response to declines in the spot and forward U.S. Treasury 10-year rates, nonbank lenders, such as investment funds, structured investment vehicles, and other shadow bank entities, acquire riskier term loans and banks facilitate these acquisitions.[4] Their results also indicate that nonbank lenders tend to acquire much riskier loan portfolios of term loans than banks do.[5] To address potential endogeneity of credit risk and interest rates, they exploit the panel structure of their data and also rely on a sample of loans made to non-U.S. borrowers. Other papers providing evidence nonbanks' risk taking following monetary accommodation include Chodorow-Reich (2014) and Maggio and Kacperczyk (2016), but they both focus on the U.S. financial market only.

In contrast, our main focus is to study the global lending channel and the cross-border spillovers of risk taking. To that end, we use a commercial data set with extensive coverage of cross-border loans— term loans made by U.S. and non-U.S. lenders to U.S. and non-U.S. borrowers in the global syndicated market—over nearly two decades. The majority of borrowers are leveraged and borrow substantial amounts in the market. Because of the richness of our data, we can differentiate lenders and borrowers by their type, nationality, and credit ratings. In addition, the longer series of the data allows us to compare the different risk-taking channels of monetary policy before and after the crisis.

# 3 Data

Our data on syndicated loans are from Thomson Reuters Loan Pricing Corporation (LPC)'s DealScan, covering the period between 1995 and 2013. This dataset contains detailed information on individual syndicated loans, borrowers' characteristics, and composition of the lending syndicates, which allows an in-depth analysis of the riskiness of the

[4]They also examine risk taking in response to an increase in the expected extent and duration of the zero lower bound period.

[5]Aramonte, Lee, and Stebunovs (2015) measure a loan's ex ante credit risk with a Basel II-consistent probability of default at a one-year horizon as reported by agent banks to their supervisors. Because banks report institutional-level ownership shares in loan syndications, they can construct acquisitions and holdings of syndicated loans by thousands of bank and nonbank lenders and assess their riskiness.

loans.

### 3.1 Global syndicated loan market

A syndicated loan is different from a simple bank loan or a corporate bond. It is extended to a borrower by multiple lenders which form a syndicate for that purpose, and it is administered by an agent, typically a bank. The loan's interest rate is floating; it is equal to a contractual loan spread over a benchmark interest rate. The spread is determined at origination, reflects lenders' judgments about the loan's credit risk, and stays constant over the loan's duration unless the loan's covenants are violated. For a given loan, its interest rate changes as the benchmark interest rate moves, therefore, minimizing lenders' interest rate risk.

In our analysis, we focus on term loans that are denominated in U.S. dollars, indexed to the U.S. dollar LIBOR, and originated in the global syndicated loan market. These loans are dispersed at origination, usually have a maturity of 5-7 years, and therefore, are more similar to corporate bonds than credit lines. We exclude credit lines in the analysis for two reasons. First, while a variety of different types of lenders acquire term loans, the lenders of credit lines are predominantly banks (see the findings in Aramonte, Lee, and Stebunovs (2015)). Second, credit lines usually have complex pricing and drawdowns, which are more likely to be endogenous to the credit cycle and their information are not available in our dataset.

In terms of market size, originations of syndicated term loans, are comparable to issuance of nonfinancial corporate bonds. See panel (a) in Figure 1. Starting from early 2000s, new loan originations had been rising at a fast pace through 2007 but subsequently collapsed during the crisis. Since reaching the trough in 2009, the syndicated loan markets had recovered substantially. Originations of term loans reached to about \$2 trillion in 2013, of which roughly \$700 billion made to U.S. borrowers and \$1.3 trillion to non-U.S. borrowers. Syndicated lines of credit, extended mostly to borrowers in advanced economies, added another \$2 trillion to originations in the overall syndicated loan markets.

To determine the identity of lenders and borrowers for each individual loan, we analyze lending on an ultimate counterparty basis. Specifically, we aggregate lenders up to their parent organizations—ultimate lenders. Similarly, we use locations of parent organizations—ultimate borrowers—to determine locations of syndicated term loan borrowers. The approach reflects the notions that risk management in financial institutions is carried out at a parent-company level, and that borrower funding decisions are also made at a parent-company level. In addition, parent companies tend to guarantee debt being borrowed by their subsidiaries.

For identification purposes, we also need to distinguish lenders and borrowers by their types. We rely on two-digit SIC codes for primary activity given in the DealScan data to classify lenders. We consider two types of lenders: banks and nonbank financial institutions, which include finance companies; investment banks; investment funds; structured products, and insurance companies. As demonstrated below, these nonbank financial lenders have modest ownership shares in the origination of syndicated term loans, thus economically it is more meaningful to group them together rather than analyze them separately.[6] For borrowers, we rely on one-digit SIC codes to identify primary industries of borrowers.[7]

As noted above, one of the advantage of studying the global syndicated loan market is its wide coverage of lenders of various types and nationalities supplying credit to borrowers of varying credit quality from around the world. In terms of risk profile, a vast majority of syndicated term loans are extended to risky, opaque borrowers. As illustrated in Figure 1 panel (b), loans to nonrated borrowers dominated originations, whereas loans to investment-grade and speculative-grade borrowers accounted for only a moderate fraction. Loan spreads on loans to nonrated borrowers are typically much higher than those on loans to investment-grade borrowers, but somewhat lower than those on loans to speculative-grade borrowers. In our data, risky loans—with contractual loan spread over 125-150 points—account for

[6] We can perform the analysis by lender type at a finer level and show that the results hold for a broad set of nonbank lenders, but economic significance of these results may be lacking because of a small number of lenders in each of the finer lender types.

[7] In a limited number of cases when primary industry identifiers are missing for parent companies, we rely on primary industry identifiers of their subsidiaries.

the bulk of loan originations. Loans extended to nonrated and speculative-grade borrowers tend to be leveraged. In terms of geographical composition, loans extended to borrowers from EMEs accounted for a sizable share at origination. Panel (c) in Figure 1 and panel (c) in Figure 2 illustrate the volume composition and pricing of syndicated term loans extended to different groups of borrowers.

From the lenders side, nonbank lenders tend to originate riskier loans than banks do, as the contractual loan spreads for new loans made by nonbanks are significantly wider than the spreads for loans made by banks.[8] In terms of market share, banks and bank holding companies (BHCs) accounted for about 80 percent of originations volume globally, and the rest was originated by nonbank lenders.[9] See panel (a) in Figure 3. The dominance of banks and BHCs at origination is partially driven by the fact that banks and BHCs often have a better lending network and stronger capacity to evaluate the credit risk of borrowers comparing to nonbanks.

However, according to a separate date source—the confidential supervisory Shared National Credit (SNC) data, banks often originated syndicated loans with intent to distribute to lenders of other types that either pre-committed to buying shares in the loans being originated or likely to buy shortly after the origination. Indeed, banks' ownership shares in seasoned syndicated term loans tend to be much lower than those at origination. In the U.S. market, banks and BHCs own only over a third of all term loans. Nonbank lenders such as investment funds, structured finance products, portfolio managers, finance companies, and other types account for the rest. From this perspective, nonbanks are well represented in the syndicated term loan market even though their share at origination does not appear to be large. To illustrate how fast banks sell their loan stakes to nonbank lenders, Figure 3 panel (b) plots the dynamics of banks' ownership share during the period after after loan originations in the U.S. market. As shown in the picture, the median of banks' ownership shares of syndicated loans declines from about 80 percent to less than 20 percent in just a

[8]Consistently, as Aramonte, Lee, and Stebunovs (2015) find, the default probabilities of loans held by nonbank lenders (as identified in the SNC data) were also much higher than those held by banks.

[9]Moreover, among banks, originations were dominated by a group of select lenders.

few weeks after the origination.

### 3.2 U.S. interest rates

U.S. interest rates are another variable that is of main interest in our analysis. We use effective federal funds rate as a measure of short-term interest rate and 10-year U.S. Treasury (zero-coupon) rate as a measure of long-term rate. The former is the Federal Reserve's key monetary policy target in the pre-crisis period, while the latter is an implicit target in the post-crisis period after the federal funds rate reached the zero lower bound. Panel (a) in Figure 2 plots the time series of the two interest rates in our sample period. In the pre-crisis period, both short- and longer-term U.S. interest rates varied in a wide range. The two rates moved closely in the second half of 1990s, but subsequently diverged until 2006. The downward trend in 10-year U.S. Treasury rate over this period was possibly driven by better-anchored inflation expectations and the savings glut—two factors not related to the credit cycle. In the post-crisis period, though short-term interest rates stayed near zero, the 10-year U.S. Treasury rates declined initially but rebounded moderately in 2012-13.

### 3.3 Loan spreads and ex-ante credit risk

Before presenting our main regression analysis, we first demonstrate that loan spreads in the primary syndicated loan market are reliable proxies for borrowers' ex ante credit risk, measured in terms of probabilities of borrower defaults. Several previous studies established the link between loan spreads and credit risk. For instance, Strahan (1999) argues that both the price and non-price terms of bank loans reflect observable components of borrower risk. Using Dealscan data, he shows that riskier borrowers—smaller borrowers, borrowers with less cash, and borrowers that are harder for outside investors to value—pay more for their loans. In addition, the non-price terms of loans are systematically related to pricing: Small loans, secured loans, and short-maturity loans carry higher interest rates than other loans, even after controlling for publicly available measures of risk.[10] In more recent work, Gaul

[10]This suggests that banks use both the price and non-price terms of loans as complements in dealing with borrower risk.

(2014) shows that loan spreads at origination forecast future loan downgrades from pass to substandard and doubtful ratings, and that their predictive power is somewhat higher than that of a set of publicly available metrics of borrowers' credit risk.[11]

The theoretical underpinning of loan spreads as proxies for ex-ante probability default can be illustrated in a basic loan pricing model. For example, consider a risk-neutral perfectly competitive lender prices a loan which, for simplicity, has no guarantees or collateral. In this simple setup, perfect competition in the loan market and zero profit condition imply that the loan spread is proportional to the product of the probability of loan default and loss given default rate.[12] In practice, however, syndicated term loans tend to have maturities of several years, many of these loans may have guarantees, collateral requirements, maintenance covenants, and some of them may be not senior, so the pricing of syndicated loans may be more complicated than the model suggest.[13] Nevertheless, as we will show below, loan spreads still reflect a large share of variations in the ex-ante probability of loan default cumulated (averaged) over the entire duration of the loan.

To test the empirical relationship between loan spreads and credit risk, we follow Aramonte, Lee, and Stebunovs (2015) and merge the commercial DealScan and quarterly SNC data together by loan terms (hard information). We then verify the matches using borrower names, which may be spelled differently in these two data sets (soft information). Because of the scope of the SNC data, our sample is restricted to the U.S. syndicated loan market, where both U.S. and non-U.S. lenders provide credit to U.S. borrowers and some non-U.S. borrowers. The sample covers the period from 2010 to 2013, as the quarterly data collection for the SNC program began only in 2009:Q4.

[11] Gaul and Stebunovs (2009) and many others present evidence that loan spreads also reflect private information about borrowers' risk characteristics. In addition, other market dynamics, market microstructure, overhead costs may contribute to the final determination of spreads. We will examine how such factors determine spreads later on in this section.

[12] The model can be also extended to include risk-averse lenders with pricing power and financing constraints of their own. In the more elaborated setup, other factors, such lenders' risk aversion, pricing power, and cost of funds (in excess of an increase in the U.S. dollar LIBOR, the benchmark rate used in pricing), could also determine loan spreads.

[13] Some loans could also get refinanced and the new loan may have a different spread. In addition, if the quality of the loan deteriorates over time, the spread may increase in a pre-determined fashion, as the syndicated loan contract may stipulate.

We measure ex-ante credit risk using a Basel II-consistent risk parameter—through-the-cycle probability of borrower default at a one year horizon (PD).[14] Figure 4 plots a simple correlation between loan spreads (all-in-drawn spreads) and probabilities of loan default, which shows that ex-ante riskier loans command higher loan spreads and that the relationship between the two is highly nonlinear (note that both PDs and loan spreads on the axes are logged).[15] Table 1 reports some distribution statistics for the data presented in the figure. In line with our main dataset of the long-series global syndicated term loans, the matched DealScan-SNC data sample is also dominated by term loans to non-rated borrowers. As noted above, these borrowers typically have probabilities of default much higher than those of investment-grade borrowers but somewhat lower than those of speculative-grade borrowers.[16]

To demonstrate more formally the relationship between loan spreads and ex-ante credit risk, we estimate the following regression:

$$\begin{aligned} log(spread_{j,l,b,t}) = \alpha_l + \delta_{PD} log(PD_{j,b,t}) + Q_{j,t}\delta_Q \\ + \beta_T R_t^T + X_t\gamma + \theta_{l,y} + \theta_{b,i,y} + \varepsilon_{j,l,b,t} \quad (1) \end{aligned}$$

where $spread_{j,l,b,t}$ is the all-in-drawn spread for loan $j$ made/reported by lender $l$ to borrower $b$ at time $t$; $PD_{j,b,t}$ is the probability of default of borrower $b$ that received loan $j$; $Q_{j,t}$ is the vector of loan $j$'s characteristics such as loan maturity and purpose; $R_t^T$ is the 10-year U.S. Treasury (zero-coupon) interest rate; $X_t$ is the vector of controls for risk appetite and

[14] The coverage of other risk parameters is rather limited in the SNC data, so we limit our analysis to probabilities of default. As in Aramonte, Lee, and Stebunovs (2015), we capture the ex-ante credit risk of each loan with the default probability that the reporting bank uses to determine regulatory capital. Regulations require banks to use through-the-cycle default probabilities that provide an assessment of a loan's credit risk at a one-year horizon.

[15] The all-in-drawn spread include annual fees, prorated based on loan maturity. In our sample, only a very small portion of loans have annual fees. Note that term loans are consistently drawn in full at origination. They have fewer built-in optionalities than credit lines and, hence, have fewer types of fees. As Berg, Saunders, and Steffen (2016) point out, the upfront fee is by far the most frequent of term loan fees, occurring nearly 30 percent of the time.

[16] As Aramonte, Lee, and Stebunovs (2015) point out, ratings appear to play a significant role in pricing of loans. (Ratings are based on Moody's senior debt ratings at the moment of origination as reported in DealScan.) Borrowers with outstanding debt rated as investment grade generally pay significantly lower loan spreads than borrowers with either not rated or speculative grade debt.

macroeconomic factors; $\theta_{l,y}$ and $\theta_{b,i,y}$ are reporting bank $l$-year fixed effect and borrower industry-year fixed effects respectively. These fixed effects control for lender-specific and borrower industry-specific factors that change at an annual frequency. The regression also includes lender $j$'s fixed effect $\alpha_l$ but not borrower $b$'s fixed effect because we have too few repeated borrowers in the sample. $\varepsilon_{j,l,b,t}$ is a white noise error. We cluster errors by time because multiple loans originated in a given period are subject to the same macroeconomic and regulatory environment.[17]

The estimation results are reported in Table 2. As shown in column (1), PDs appear to play an important role in pricing syndicated loans. There is a positive, statistically significant relationship between log(spread)s and log(PD)s; PDs explain about 40 percent of variation in loan spreads. In column (2), we add loan characteristics which pushes up the explanatory power by 10 percentage points. That is, inclusion of loan characteristics, which may be correlated with probabilities of default, improves the goodness of fit by several percentage points but does not materially affected the explanatory power of probabilities of default.

In column (3), we add various macroeconomic and financial variables that can potentially affect spreads—for example, risk aversion (a VIX-derived variance risk premium and spreads of U.S. low-grade corporate bonds over the U.S. Treasury rate) and reporting bank-year fixed effects to account for agent characteristics that change gradually (for example, its capital and liquidity positions).[18] Adding these variables improves the adjusted R-squared by only 6 percentage points, mostly due to reporting bank-year fixed effects, as regression coefficients on variance risk premium and other risk appetite variables are statistically insignificant. These results suggest that log(PD)s explain a significant portion of variations in log(spread)s, but financial and risk-aversion variables do not contribute much. In other

[17] As shown in Petersen (2009), in the presence of a time effect, standard errors clustered by time produce unbiased standard errors and correctly sized confidence intervals. The results are robust to clustering errors by either reporting bank or borrower industry.

[18] VRP here is calculated from options prices as the difference between the expected realized variance and the squared VIX index, considered to be the most readily available proxy for fluctuations in investors' risk aversion. Low-grade corporate bond spreads are from Merrill Lynch, which include corporate bonds of all maturities and all industries and are available from the early 1980s. Because low-grade bond is very risky (riskier than high-yield bonds), yields on these bonds are a good gauge of investor appetite for risky assets.

words, loan spreads reflect mostly loan risk characteristics rather than variation in degrees of lenders risk aversion. In column (4), we add borrower industry-year fixed effects to account for borrower industry characteristics—for example, demand for bank loans—that change gradually. These additional fixed effects again improve the adjusted R-squared only marginally.

To sum up, the empirical results above suggest that loan spreads are a reliable proxy for ex-ante credit risk and risk aversion plays only a minor role in pricing syndicated term loans. Depending on the scope of the analysis, to some extent, loan spreads may be seen as more comprehensive measures of ex-ante credit risk than PDs alone, as they also reflect other risk characteristics of loans.[19]

# 4 Empirical methodology

We use the panel structure of the data to help with identification of risk-taking in response to interest rates. Under this structure, although not perfect, one can assume that each lender takes U.S. short-term and long-term interest rates as given and makes their portfolio decisions in response to changes in U.S. dollar-denominated cost of funds and returns on safe assets. The identification is the strongest when we restrict our panel data to the sample of non-U.S. lenders extending loans to non-U.S. borrowers, as changes in U.S. interest rates are likely exogenous to changes in ex-ante credit risk in these loans. In addition, the panel structure allows us to introduce various fixed effects to control for unobserved factors that may affect lenders and borrowers, which helps further strengthen the identification.

Note that changes in interest rates may not only cause lenders to switch between borrowers with different risks and hence different spreads, they may also cause changes in the

[19] Because loan spreads reflect PDs and other risk characteristics, there may be a concern that our approximation of loss rates with loan spreads introduces a measurement error to our left-hand side variable in our benchmark model below. As Hausman (2001) points out, econometrically, this may not be an issue. In regressions where the right-hand side variable is measured without error, the consequences of a mismeasured left-hand side variable are innocuous. In this case, the ordinary least squares estimator would be unbiased under a wide range of assumptions, but with reduced precision in the estimate, a lower t-statistic, and a reduced R-squared.

spreads themselves, for given levels of credit risk. Hence, our measure of credit risk is subject to a problem not present with other risk proxies, such as default-probability measures. To address this possibility, we include spreads of U.S. low-grade corporate bonds over the U.S. Treasury rate in the regressions, to control for general movements in risk aversion. As the earlier results for the loan spread regressions suggest, however, the effects of premiums related to risk aversion are likely rather small.

Our benchmark models estimate the relationship between interest rates and risk taking at both the syndicate loan level and lender portfolio level, through which we examine risk taking behavior of financial institutions at different decision making stages. While the portfolio regression may reflect lenders' decision making on riskiness more realistically, the syndicate model allows for a better control for unobserved push and pull factors.

### 4.1 Syndicate regressions

Because of the skewness of the distribution of the explained variable, the benchmark syndicate model takes a semi-log form:

$$log(spread_{j,l,b,t}) = \alpha_l + \alpha_b + \underbrace{\beta_F R_t^F + \beta_T R_t^T + X_t \gamma}_{h.f.\ push\ factors} + \underbrace{\theta_{l,c,y} + \theta_{l,i,y}}_{l.f.\ push\ factors} + \underbrace{\theta_{b,c,y} + \theta_{b,i,y}}_{l.f.\ pull\ factors} + \phi_{j,t} + \varepsilon_{j,l,b,t} \quad (2)$$

where $spread_{j,l,b,t}$ is the all-in-drawn spread for loan (syndicate) $j$ made/reported by lender $l$ to borrower $b$ at time $t$; $R_t^F$ is the federal funds rate; $R_t^T$ is the 10-year U.S. Treasury (zero-coupon) interest rate; $Xt$ is the vector of controls for risk appetite and macroeconomic factors. The $X$ controls include the following variables. We use the Merrill Lynch low-grade corporate bond spread to measure risk aversion and credit risk compensation in the secondary corporate bond market, and use the VIX-derived variance risk premium, the European sovereign crisis (a spread between Italian and German sovereign yields), economic uncertainty (the U.S.-news based uncertainty index of Baker, Bloom, and Davis (2015)) to gauge lender risk appetite.[20] In addition, we include expected inflation (Michigan consumer survey-based expected 12-month inflation) to capture the notion that it is real interest rate that ultimately determine lenders' risk taking.[21]

$\theta_{l,c,y}$ and $\theta_{l,i,y}$ are lender $l$'s country group-year and type-year fixed effects respectively. In some regressions, we assign lender countries to 3 groups—the United States, non-U.S. advanced economies, and EMEs) and in others to just 2 groups (non-U.S. advanced economies and EMEs). Based on lenders' industries, we assign lenders to 3 types—banks, nonbank financial lenders, and nonfinancial lenders. These fixed effects controls for unobserved factors that affect lender behavior at an annual frequency, which could be interpreted as push factors for the supply of credit to borrowers in various industries and countries.

Similarly, $\theta_{b,c,y}$ and $\theta_{b,i,y}$ are borrower country group-year and industry-year fixed effects, respectively, which control for borrower country- and industry-specific factors that change at an annual frequency. In the regressions, we assign borrower countries to only 2 groups (non-U.S. advanced economies and EMEs) because we drop loans to U.S. borrowers from the sample to ensure exogeneity of credit risk to U.S. interest rates. We create borrower industry-year fixed effects that are based on borrower industry SIC codes at one digit level.[22] These fixed effects capture low frequency pull factors for the demand for credit by borrowers in various industries and countries. The regression also includes lender $l$'s fixed effect $\alpha_l$, borrower $b$'s fixed effect $\alpha_b$, and variables that capture syndicate characteristics $\phi_{j,t}$, such as the number of lenders in a syndicated.[23]

$\varepsilon_{j,l,b,t}$ is a white noise error. We cluster errors by time because multiple loans originated

[20] While Baker, Bloom, and Davis (2015) uncertainty index is derived from U.S. news sources, it captures uncertainty about events in both the United States and abroad. They seek to capture uncertainty about economic policy decisions, their effects on the economy, and uncertainty induced by policy inaction. They base this index on search results from 10 leading U.S. newspapers that tend to cover domestic and international developments.

[21] We use the survey-based expected inflation because the TIPS-based inflation compensation is not available prior to early 2000s.

[22] While we can define such fixed effects at a finer industry level, they may not be estimated consistently across years. In some years, we may have too few observations for borrowers of certain industries in a certain region.

[23] For larger loans, the number of lenders tends to be higher. Nonbank financial lenders may prefer to invest into somewhat different terms loans than banks, therefore, controlling for nonbank participation may be necessary. See, for example, Aramonte, Lee, and Stebunovs (2015).

in a given period are subject to the same macroeconomic and regulatory environment. Although loan data are available at monthly frequency, we estimate the model at a quarterly frequency so that, for many given lenders and borrower-industry-region groupings, loan originations reach economically significant volumes. In addition, as loan syndication takes some time (30-60 days) and the negotiation period for each loan within in a given quarter is not known, analysis at a quarterly frequency appears to be the most appropriate.[24]

The interpretation of $\beta_F$ and $\beta_T$, the main coefficients of interest, is straightforward: $\beta$s capture the sensitivity of risk taking to changes in U.S. interest rates. The hypothesis is that the sensitivity of risk taking is negative, indicating lenders load on risk in response to a decline in U.S. interest rates. Note that because the regression model takes a semi-log form, for a given loan $j$, the marginal effect around a given level of the loan spreads is:

$$\Delta(spread_{j,\ldots}) = \widehat{spread}_{j,\ldots}\beta_i \Delta R^{i,} \quad i = F, T. \tag{3}$$

In the regression output that follows, we focus on two separate risk taking channels of monetary policy: one that is associated with movements the federal funds rate and the other with the 10-year constant maturity U.S. Treasury rate. An alternative specification is to regress log(spread) on a slope of the yield curve (defined as the spread of the 10-year rate over the federal funds rate) rather than its components, but we do not do so for two reasons. First, the inclusion of the slope is equivalent to imposing a constraint on the coefficients of its components: The coefficients are restricted to be of the same absolute magnitude but of opposite signs. We prefer the data speaking about such coefficient restrictions. Second, in the post-crisis period, because the federal funds rate is at the zero lower bound, the movement in the slope would be solely driven by the 10-year Treasury rate.

In our baseline regressions, we first estimate equation 2 for the pre- and post-crisis period separately, given the structural break in the interest rate series. For the pre-crisis period, we include both short-term and long-term interest rates to test whether the two

[24]Analyzing the data at a quarterly rather than monthly frequency also helps to minimize seasonality concerns (more originations tend to be done towards quarter ends).

risk taking channels of monetary policy operate simultaneously. For the post-crisis period, because the federal funds rate was at the zero lower bound beginning 2008:Q4, the regression includes the U.S. Treasury rate only. In addition, we estimate an extended version of the regression model over the entire sample period, which includes both the federal funds rate ($R^F$) and the U.S. Treasury rate ($R^T$), as well as the interactions of these interest rates with a post-crisis dummy ($Post2008$):

$$\begin{aligned} log(spread_{j,b,l,t}) = {} & \alpha_l + \alpha_b + \beta_F R_t^F + \beta_T R_t^T \\ & + \beta_{F08} Post2008 \times R_t^F + \beta_{T08} Post2008 \times R_t^T + \beta_{08} Post2008 + \beta_C Crisis_t \\ & + X_t \gamma + \theta_{l,c,y} + \theta_{l,i,y} + \theta_{b,c,y} + \theta_{b,i,y} + \phi_{j,t} + \varepsilon_{j,b,l,t}. \end{aligned} \quad (4)$$

## 4.2 Portfolio regressions

The portfolio model captures the notion that financial institutions in general consider their risk taking strategy at the loan portfolio level, although we note that aggregation of loan syndicates limits the number of unobserved pull factors that can be included in this specification.

We calculate the average loan spread of all term loans made by each lender in each quarter. For many loans in DealScan, the detailed share of ownership in a loan is not available. To overcome this issue we either construct simple averages of loan spreads for loans where a particular lender participated, or construct weighted averages of loan spreads based on the assumption used in the previous literature that lead lenders are assigned larger shares than other lenders. The results are not affected by the averaging method. Thus, we subsequently use simple-average loan spreads and refer to them as portfolio loan spreads. Similar to the syndicate regression, our benchmark portfolio model also takes a semi-log form as follows:

$$log(pspread_{bc,l,t}) = \alpha_{bc} + \alpha_l + \underbrace{\beta_F R_t^F + \beta_T R_t^T + \gamma X_t}_{h.f.\ push\ factors} + \underbrace{\theta_{l,c,y} + \theta_{l,i,y}}_{l.f.\ push\ factors} + \underbrace{\theta_{bc,y}}_{l.f.\ pull\ factors} + \varepsilon_{l,b,t} \quad (5)$$

where $spread_{bc,l,t}$ is the average spread for loans originated by lender $l$ to borrowers in country group $bc$ at time t. $\alpha_{bc}$ is the borrower country group fixed effect. $\theta_{bc,y}$ is borrower country group-year fixed effect which represents a low frequency pull factor.[25] The other variables in the model are the same as in regression model 2.

To make the results comparable with the syndicate regressions, we first estimate the portfolio model for the pre-and post-crisis periods separately. We then re-estimate the model for the entire sample period with post-crisis interaction terms in a similar fashion as that in the syndicate regressions 4:

$$log(pspread_{bc,l,t}) = \alpha_{bc} + \alpha_l + \beta_F R_t^F + \beta_T R_t^T + \beta_{F08} Post2008 \times R_t^F + \beta_{T08} Post2008 \times R_t^T + \beta_{08} Post2008 + \beta_C Crisis_t + \gamma X_t + \theta_{l,c,y} + \theta_{l,i,y} + \theta_{bc,y} + \varepsilon_{l,b,t} \quad (6)$$

where the notation is the same as in the earlier models.

Finally, we estimate a version of regression models 5 and 6 with separate regression coefficients on U.S. interest rates for banks and nonbank financial lenders. This regression model examines whether risk taking by bank and nonbank lenders responds differently to changes in U.S. interest rates. However, we note that these estimations may not provide a complete picture of the risk taking strategy by different financial institutions. As stated above, although banks dominate loan originations, they quickly sell these loans to investors

[25] As an alternative, we also use higher-frequency pull factor, such as average real GDP growth of borrower countries weighted by loan originations. The results are stronger as the alternative specification may have missed a variety of latent pull factors.

of other types who may pre-committed to the purchase. Hence, the estimated coefficients $\beta_F$ and $\beta_T$ for banks may reflect risk taking by nonbanks to some extent.

Table 3 reports summary statistics for our main explained and explanatory variables. Since our regression analysis focus on non-U.S. borrowers, we report the loan spreads for this group and EME borrowers only. Interestingly, the average loan spread for EME borrowers at 1.67 percentage points is somewhat lower than that for non-U.S. borrowers (1.86 percentage points), which includes borrowers from both non-U.S. advanced economies and EMEs. The standard deviation is also smaller. This may partially reflect the fact that mostly large firms with decent credit from EMEs borrow in the global syndicated loan market. The average federal fund rate is 3.0 percent with a standard deviation of 2.3 percent. In contrast, the 10-year Treasury rate displays a much smaller variation over the sample period, with a standard deviation of 1.4 percent.

# 5 Estimation results

This section presents our main findings regarding the effects of changes in U.S. interest rates on risk taking in the global syndicated loan market. In particular, we differentiate the roles played by short-term and long-term interest rates in the pre- and post-crisis periods, as well as the response by bank and nonbank lenders. We also present several robustness checks to enrich our main findings and demonstrate that our baseline results are not driven by certain assumptions used in the analysis.

## 5.1 Syndicate regressions

**Loans made to non-U.S. borrowers by any lenders.** We first estimate models 2 and 4 over the sample of all non-U.S. borrowers and lenders on an ultimate counterparty basis. This is the broadest data cut that we can do while separating U.S. monetary cycle and non-U.S. credit cycles. The estimation results are reported in Table 4. Column (1) shows the results of regression model 2 for the pre-crisis period, and column (2) for the post-crisis period. In the latter specification, again the federal funds rate is excluded, because it was at the lower zero bound and its effect cannot be reliably estimated. In the pre-crisis period, the results point to a significantly negative relationship between short-term interest rates and ex-ante risk taking, while the effects of movements in long-term rate is statistically insignificant. In contrast, the post-crisis period is characterized by a significantly negative relationship between long-term interest rate and risk taking. Column (3) shows the results for regression model 4 over the entire sample period of nearly two decades, including the global financial crisis episode. The result for the federal fund rate in the pre-crisis remains similar, but the coefficient for the long-term rate is no longer statistically significant. This is likely due to the inclusion of the crisis quarters in this estimation.[26]

Based on the regression estimates in columns (1) and (2), we can calculate the marginal effects in the neighborhood of average spreads using equation 3.[27] The calculations suggest that, in the pre-crisis period, a one-standard deviation decrease in the federal funds rate of 1.82 percentage point leads to an increase in loan spread by 11 basis points. In comparison, in the post-crisis period, a one-standard deviation decrease in the 10-year Treasury rate of 0.71 percentage point raises the loan spread by 28 basis points. Hence, the economic effect of the long-term rate on risk taking in the post-crisis period appears to be much stronger than that of the short-term rate in the pre-crisis period. These results are particularly striking because the standard deviation of the longer-term interest rate is significantly lower than that of the federal funds rate. The changes in loan spread can then be further mapped into changes in loan PDs based on the estimation results of regression model 1.

The estimated coefficients for the other explanatory variables are of interest too. While the coefficient on the variance risk premium is statistically insignificant across the specifications, the negative coefficient on the uncertainty index is statistically significant in the pre-crisis period, indicating that elevated uncertainty moderated lenders' appetite for riskier loans.[28] The coefficient on low-grade bond spreads is positive and consistently statis-

[26]Later, we show that on a portfolio basis, the effects of long-term rate on risk taking is present in both the post-crisis period and in the longer period that includes the crisis.

[27]The average spread during the pre-crisis period is about 100 basis points and the average in the post-crisis period is about 260 basis points.

[28]The result that uncertainty index can outperform the VIX and VIX-derived indexes is consistent with the findings in other studies. For example, Baker, Bloom, and Davis (2015) examines the effects of uncertainty

tically significant, which implies that a compression of loan spreads could be associated with broad improvements in investors' appetite for term loans. The coefficient on the European sovereign yield spread is only negative in column (2), likely because this period captures the European sovereign debt crisis and the risk off environment associated with it.

**Loans made to non-U.S. borrowers by U.S. lenders.** Next we restrict our sample to loans made to non-U.S. firms by U.S. lenders only. This sample cut is in the spirit of the nascent global lending channel literature which explains the magnitude and direction of U.S. bank lending to international borrowers caused by changes in U.S. monetary policy. Table 5 summarizes our results. The regression results are very similar to those reported in Table 4, where we find evidence of different risk taking channels of monetary policy in the pre- and post-crisi periods. Noticeably, for the U.S. lender sample, the magnitude of the estimated coefficient for 10-year Treasury rate in the post-crisis period is much larger in absolute terms, which suggests that the economic effect of long-term rate changes on risk taking is somewhat stronger for U.S. lenders than for non-U.S. lenders. Overall, even restricting the sample to U.S. lender–non-U.S. borrower loans, we find evidence that low interest rates in the U.S. induce risk taking in the global syndicated term loan market.

**Loans made to EME borrowers by non-U.S. lenders.** The final exercise in the syndicated model is to restrict our sample to loans made to firms in EMEs only by non-U.S. lenders. In terms of exogeneity, this is the strongest specification that we have, as the credit and business cycles in EMEs are less likely to co-move with U.S. interest rates comparing to the cycles in non-U.S. advanced economies. In addition, it is also the strongest specification to demonstrate the cross-border spillover effects of monetary policy through the global lending channel. Table 6 summarizes our results. Again, most results are very close to those presented in Table 4. This suggests that the earlier results are not driven solely by the interactions of U.S. lenders and borrowers from non-U.S. advanced economies.

on firm-level investment and employment using both the VIX and the uncertainty index. They show that only the latter has negative and statistically significant effects.

In sum, our main results are robust to restricting the sample to EME borrowers only. In order words, we provide evidence that lower interest rates in the U.S. induces non-U.S. lenders to extend loans to risker borrowers from EMEs in the global syndicated term loan market.

## 5.2 Portfolio regressions

We now summarize the regression results of the portfolio models. This setup is more consistent with the idea that the lenders make decisions on risk management and risk taking based on their portfolios of loans. To strengthen the identification, we consider each lender manages two types of portfolios: one comprised of loans to borrowers in non-U.S. advanced economies and the other comprised of loans to EME borrowers. This allows us to introduce borrower country group-year fixed effects that control for credit demand and other unobserved factors.

**Loans made to non-U.S. borrowers by any lenders.** Table 7 reports the estimation results for this subsample. Similar to the results in Tables 4 and 6, we find that the federal funds rate has a negative and statistically significant effect on ex-ante risk taking in the pre-crisis period, whereas the the 10-year Treasury rate has a negative and statistically significant effect in the post-crisis period. Unlike the results in the syndicated loan regressions, however, the coefficient for the interaction term of 10-year Treasury rate and the post-crisis dummy is negative and statistically significant in the full sample specification. That is, our full-sample portfolio model confirms the differential effects of short-term and long-term interest rates on ex-ante risk taking in the pre- and post-crisis periods. Indeed, the difference in magnitude between the two channels is amplified based on the larger differences between the estimated coefficients on the fed funds rate and the 10-year Treasury rate. These results reinforces the findings based on syndicated loan level regressions, and provides greater economic significance to our findings. Some other differences to be noted are that, in the portfolio level regressions, the European sovereign spread and low-grade bond spread are statistically significant only in column (3) of the full sample specification.

**loans made to non-U.S. borrowers by any lenders of two types—banks and nonbanks** Finally, we test the hypothesis on whether risk taking by bank and nonbank lenders responds differently to changes in U.S. interest rates. Table 8 reports the results. The estimation results indicate that banks adjust their risk portfolios in response to changes in the federal funds rate before the crisis and also to the 10-year Treasury rate after the crisis. In contrast, nonbank lenders only respond to changes in the 10-year Treasury rate in the post-crisis period. These findings to some extent are in line with the implications from the theoretical literature, in the sense that risk taking by banks is likely driven by both the cost-of-fund and return-on-safer-asset concerns, whereas risk taking by nonbank institutions are likely motivated by return-on-safer asset or search-for-yield incentives. The different responses in risk taking by bank and nonbank lenders are particularly interesting given the gradual increase in the share of nonbank participants in the global syndicated term loan market. As such, it may present different policy implications in addressing the financial stability concern related to risk taking.

## 5.3 Robustness checks and caveats

In the appendix, we offer a few robustness checks. First, we note that while lenders and borrowers' risk management decisions are generally made at a parent level, some decisions to lend or borrow may be made on a local or immediate basis. Hence, we reestimate regression models 2 to 6 based on an immediate counterparty basis. The results on an immediate basis are weaker but still consistent with those on an ultimate counterparty basis. Second, we estimate our models using two additional subsamples—loans extended to nonrated and speculative borrowers (i.e. excluding investment-grade borrowers) and loans originated by the most active lenders (lenders that lend in the market in every quarter). We find that omitting loans to investment-grade borrowers and restricting loans from the most active lenders give statistically stronger results.

Finally, we acknowledge that our work has a couple of caveats. First, we do not observe lenders' entire portfolios, which may include a variety of instruments. For example, lenders may hedge exposures to the syndicated term loan market in other financial markets. In this case, additional risk taking in the syndicated loan market as a result of low interest rates may not necessarily represent a general increase in the riskiness of the financial sector. Second, as noted above, our results on the risk taking channel through long-term interest rates in the post-crisis period may be partly due to expansionary monetary policies adopted in other large advanced economies, not just in the United States. Thus, these results may be better interpreted as the international spillovers of risk taking associated with low interest rates in general rather than U.S. monetary policy per se.

## 6 Conclusions

This paper provides robust evidence on the negative relationship between interest rates and ex-ante risk taking in the global syndicated term loan market. Our empirical analysis differentiates the risk taking channels associated with short-term and long-term interest rates before and after the global financial crisis. We find that, in the pre-crisis period, as U.S. short-term interest rate declined, all else held equal, lenders of various types and nationalities tended to originate riskier loans to both U.S. and non-U.S. borrowers. In the post-crisis period, however, lenders of various types and nationalities tended to originate riskier loans in response to declines in long-term interest rates. The results are particularly stronger when we restrict our sample to non-U.S. lenders and borrowers from EMEs.

Our findings also present evidence of cross-border spillovers with low interest rates in one country encouraging risk taking in other countries. Such spillovers highlight the potential challenges faced by central banks, particularly those in EMEs, in affecting risk taking in lending and, more broadly, credit cycles in their respective jurisdictions. For example, while a central bank is tightening to cool off the country's credit cycle, other foreign central banks may be loosening, and, hence, inadvertently encouraging internationally active lenders to extend credits to riskier borrowers in that country. To this end, central banks may have to rely on coordination of multiple policy tools, such as monetary and macro-prudential policies, to ensure both economic and financial stability (see Rey (2015) and Obstfeld (2015) for further policy discussions related to the global financial cycle).

Our analysis contribute to the recent policy debate on whether the prolonged low interest rate environment has caused increasing vulnerabilities in the financial system. As noted above, our evidence on additional risk taking is limited to the global syndicated loan market only and lenders may have hedged potential losses from syndicated loan defaults. Nevertheless, through the hedge, other financial intermediaries may have to compensate for these losses if defaults were to occur. Hence, in aggregate, the financial system could be strained as a result of additional risk taking. That being said, increases in risk taking in the aftermath of the crisis may have also aided economic recovery. Therefore, similar to other empirical studies in this literature, our results do not address the question on whether the additional risk taking induced by expansionary monetary policy is excessive or not. This would require a comprehensive welfare analysis.

In sum, an increase in risk taking by lenders is one of the channels by which accommodative monetary policy is intended to spur economic activity. At the same time, of course, greater risk taking may pose risks to financial stability domestically and globally. Assessing the trade-off between these effects is beyond the scope of this paper, but is clearly a high priority both at the Federal Reserve and at central banks around the world.

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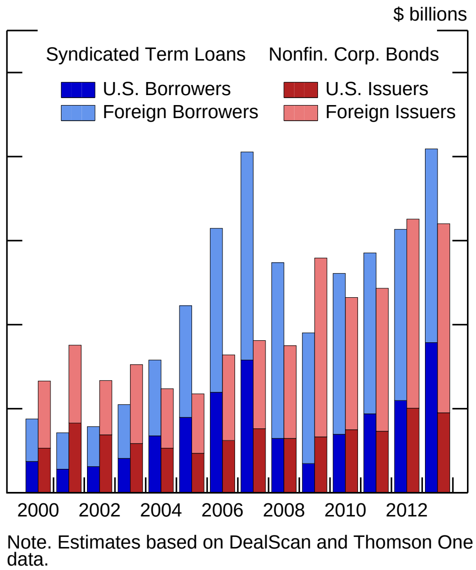
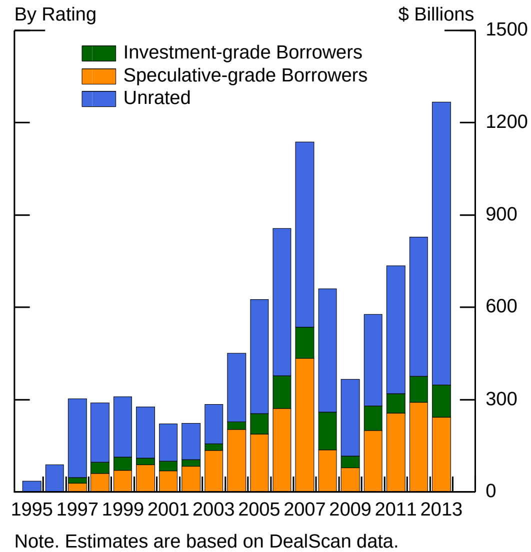
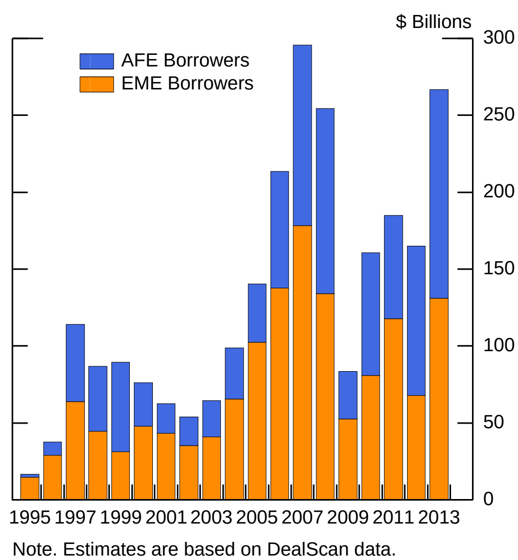



(a) Issuance of bonds and originations of loans



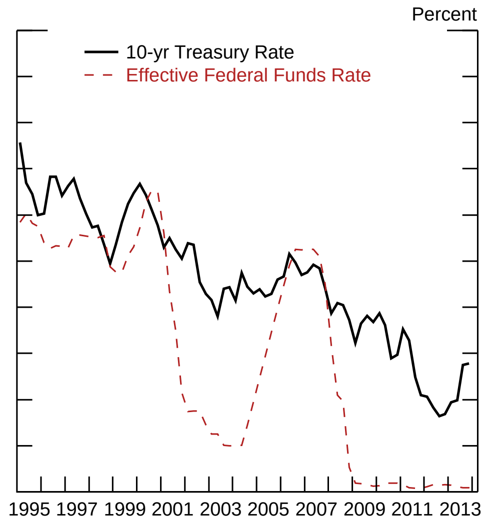

(b) Originations of loans by borrower debt rating



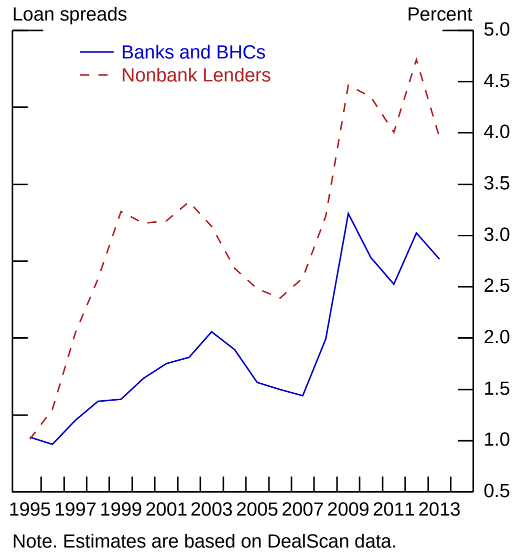

(c) Originations of loans by borrower country

Figure 1: **Issuance of corporate bonds and origination of syndicated term loans**

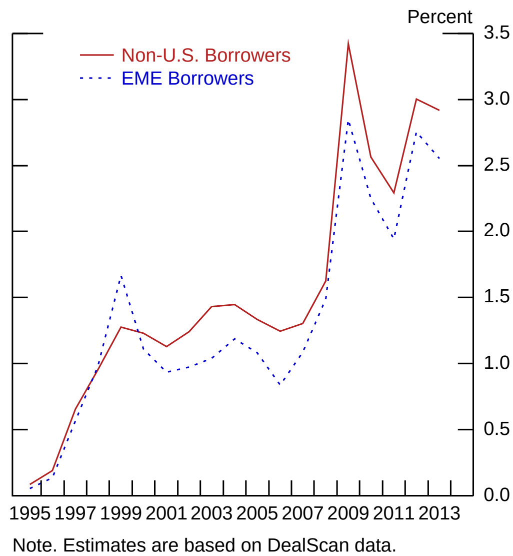
The figure is based on syndicated term loans that are denominated in U.S. dollars, indexed to the U.S. dollar LIBOR, and originated in the global market. In panel (a), originations of syndicated term loans are comparable in size to issuance of nonfinancial corporate bonds. In panel (b), the majority of syndicated term loans are made to speculative-grade and unrated debt borrowers; these loans tend to be leveraged. In panel (c), originations of syndicated term loans made to non-U.S. borrowers that are used to estimate regression models 2 to 6.




(a) U.S. interest rates



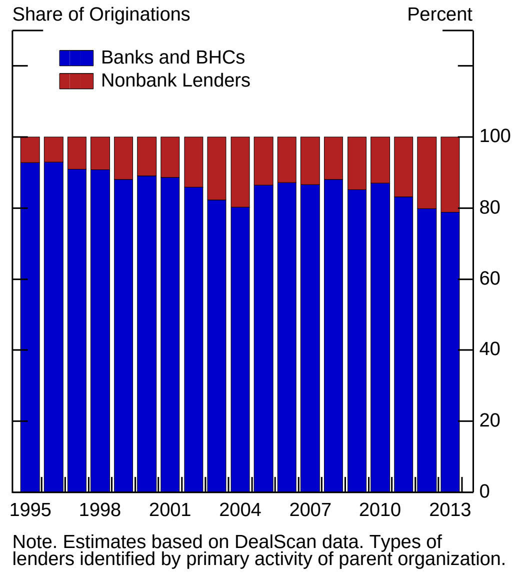

(b) Loan spreads by lender type



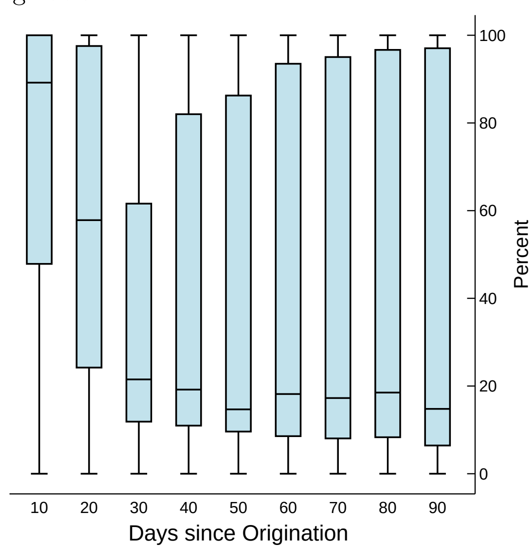

(c) Loan spreads by borrower country

Figure 2: **U.S. interest rates and pricing of syndicated term loans**

Panel (a) suggests breaking up the sample into pre- and post-crisis periods as the federal funds rate reached the zero lower bound in late 2008. Panels (b) and (c) are based on syndicated term loans that are denominated in U.S. dollars, indexed to the U.S. dollar LIBOR, and originated in the global market. Loan spreads are weighted by loan amounts. In panel (b), nonbank financial lenders tend to lend to borrowers of higher credit risk as suggested by higher loan spreads. In panel (c), loans made to EME borrowers tend to be somewhat less risky than loans made to AFE borrowers.



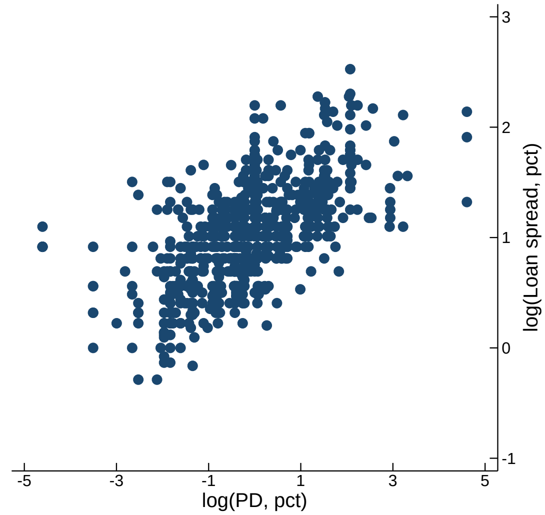

(a) Banks' ownership share of syndicated loans at origination




(b) Banks' ownership share of syndicated loans since origination

Figure 3: **Ownership of syndicated term loans at origination and over time**

Panel (a) is based on syndicated term loans that are denominated in U.S. dollars, indexed to the U.S. dollar LIBOR, and originated in the global market. It shows that banks now account for 80 percent of loan ownership at origination. Panel (b) is based on syndicated term loans made in the United States in the last several years. It shows a sequence of box plots for bank ownership shares at different time intervals since loan originations. The panel suggests that banks sell off their shares of loans to nonbank financial intermediaries—such shadow bank entities as funds and structured products—quickly. In fact, the median of bank shares in syndicated loans falls from 90 percent at origination to 20 percent within a month. Similar data for the global market are believed not to exist.



Note. Based on merged SNC and Dealscan data.

Figure 4: Loan spreads and probabilities of borrower default

The figure is based on syndicated term loans made in the United States in the last several years. It shows that loan spreads that are fixed at origination for the duration of loans are highly positively correlated with through-the-cycle probabilities of borrower default over a one year horizon.

Table 1: Descriptive statistics for loan pricing regressions

| Variable | *Mean* | *Std.dev.* | 10*th p.* | 50*th p.* | 90*th p.* |
|---|---|---|---|---|---|
| Loan spread, pct | 3.101 | 1.633 | 1.500 | 2.750 | 5.000 |
| Probability of default, pct | 2.317 | 7.110 | 0.170 | 0.800 | 4.820 |
| 10-year Treas. rate, pct | 2.346 | 0.581 | 1.710 | 2.050 | 3.200 |
| Variance risk premium, pct sq. | 14.439 | 6.218 | 7.365 | 14.034 | 25.099 |
| Sovereign yield spread, pct | 2.954 | 1.005 | 1.597 | 2.867 | 4.367 |
| Low-grade bond spread, pct | 4.855 | 0.884 | 3.622 | 4.840 | 5.839 |
| High-yield CDX, pct | 4.873 | 1.033 | 3.482 | 4.819 | 6.300 |
| Expected inflation, pct | 3.222 | 0.341 | 2.900 | 3.100 | 3.800 |

Note. The sample includes loan originations with available probabilities of default in the Shared National Credit data that are matched with loan originations in the DealScan data. The sample period is 2010:Q1-2013:Q4 because of the limited availability of probabilities of default. Loan spread statistics based on over 700 term loans, denominated in U.S. dollars, index to the U.S. dollar LIBOR, and originated in the U.S. syndicated loan market. Other variables statistics based on 16 quarters of data.

Table 2: Loan spreads as proxies for ex-ante credit risk

| | (1) PD | (2) Loan char. | (3) Macro/Bank FE | (4) Bank FE/Borr. FE |
|---|---|---|---|---|
| log(PD, pct) | 0.231*** | 0.206*** | 0.190*** | 0.189*** |
| | (10.773) | (12.069) | (10.489) | (9.870) |
| log(loan amount, $ bill.) | | −0.030** | −0.036*** | −0.040*** |
| | | (−2.411) | (−2.987) | (−3.551) |
| log(duration, years) | | 0.166*** | 0.128*** | 0.157*** |
| | | (5.819) | (3.907) | (4.116) |
| LBO or takeover loan purpose | | 0.223*** | 0.151** | 0.152** |
| | | (3.721) | (2.496) | (2.430) |
| Nontraded loan | | −0.066 | −0.120*** | −0.122*** |
| | | (−1.668) | (−5.900) | (−5.638) |
| Public company | | −0.195*** | −0.178*** | −0.175*** |
| | | (−6.351) | (−6.301) | (−6.124) |
| Non-IG borrower rating | | 0.194*** | 0.193*** | 0.179** |
| | | (3.042) | (3.106) | (2.841) |
| 10-year Treas. rate, pct | | | −0.072 | −0.093 |
| | | | (−0.964) | (−1.367) |
| VRP, pct sq. | | | 0.005 | 0.004 |
| | | | (1.197) | (1.225) |
| European sovereign spread, pct | | | 0.000 | −0.004 |
| | | | (0.010) | (−0.159) |
| Low-grade bond spread, pct | | | 0.025 | −0.017 |
| | | | (0.196) | (−0.139) |
| CDX high yield, pct | | | −0.088 | −0.052 |
| | | | (−1.074) | (−0.653) |
| Expected inflation, pct | | | 0.004 | −0.016 |
| | | | (0.073) | (−0.283) |
| News-based uncertainty index | | | −0.001 | −0.001 |
| | | | (−1.165) | (−1.420) |
| Reporting bank-year fixed effects | *No* | *No* | *Yes* | *Yes* |
| Borrower industry-year fixed effects | *No* | *No* | *No* | *Yes* |
| Num. of observations | 709 | 709 | 709 | 709 |
| R-sq. adj. | 0.40 | 0.50 | 0.56 | 0.57 |
| RMSE | 0.37 | 0.34 | 0.32 | 0.31 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.
Note. The output shown is for regression model 1. Based on DealScan loans matched with the Shared National Program data, which primarily cover loans to U.S. borrowers in the U.S. market. Dependent variable is from DealScan and independent, loan-specific variables from the Shared National Program. Dependent variable is log(loan spread, pct) of loan $j$ reported by agent bank $l$ in quarter $t$. Inclusion of loan characteristics improves the goodness of fit but does not affected the explanatory power of log(PD, pct) where PD is the through-the-cycle probability of borrower default used by the reporting bank in calculation of risk weights. Reporting bank-year fixed effects capture bank-specific conditions, such as capital and liquidity pressures, that change at an annual frequency. Borrower industry-year fixed effects capture borrower industry-specific conditions, such as creditworthiness and demand for credit, that change at an annual frequency. Fixed effects are not shown. Errors clustered by quarters.

Table 3: Descriptive statistics for loan portfolio spread regressions

| Variable | *Mean* | *Std.dev.* | 10*th p.* | 50*th p.* | 90*th p.* |
|---|---|---|---|---|---|
| Loan spread, pct | | | | | |
| Non-U.S. borrowers | 1.860 | 1.529 | 0.425 | 1.400 | 3.875 |
| EME borrowers | 1.667 | 1.352 | 0.425 | 1.250 | 3.500 |
| Fed. funds rate, pct | 3.010 | 2.336 | 0.119 | 3.056 | 5.723 |
| 10-year Treas. rate, pct | 4.474 | 1.414 | 2.414 | 4.533 | 6.354 |
| Variance risk premium, pct sq. | 19.577 | 13.423 | 6.976 | 16.149 | 34.171 |
| Sovereign yield spread, pct | 1.290 | 1.552 | 0.183 | 0.385 | 4.163 |
| Expected inflation, pct | 2.977 | 0.535 | 2.435 | 2.967 | 3.504 |
| Low-grade bond spread, pct | 5.273 | 2.550 | 3.119 | 4.705 | 7.867 |
| News-based uncertainty index | 110.801 | 39.651 | 69.307 | 100.557 | 168.622 |

Note. The sample period is 1995:Q1-2013:Q4. Loan spreads are for syndicated term loans that are denominated in U.S. dollars, indexed to the U.S. dollar LIBOR, and originated in the global market. In the sample, there are 9980 loans to non-U.S. borrowers of which 7589 are loans to EME borrowers. U.S. news source-based uncertainty index captures uncertainty in the United States and abroad, see Baker, Bloom, and Davis (2015).

Table 4: Syndicate regressions: Loans made by all lenders to non-U.S. borrowers

| | (1)<br>Pre-crisis | (2)<br>Post-crisis | (3)<br>Full |
|---|---|---|---|
| Fed. funds rate, pct | $-0.062^{**}$ | | $-0.075^{***}$ |
| | $(-2.594)$ | | $(-4.308)$ |
| 10-year Treas. rate, pct | 0.037 | $-0.150^{***}$ | $-0.010$ |
| | (0.870) | $(-3.275)$ | $(-0.253)$ |
| Variance risk premium, pct sq. | 0.002 | $-0.005$ | 0.000 |
| | (1.636) | $(-1.491)$ | (0.253) |
| European sovereign spread, pct | 0.014 | $-0.108^{***}$ | 0.003 |
| | (0.395) | $(-4.709)$ | (0.154) |
| Expected inflation, pct | $-0.036$ | $0.241^{***}$ | $0.043^{*}$ |
| | $(-1.022)$ | (3.637) | (1.707) |
| Low-grade bond spread, pct | $0.099^{**}$ | $0.127^{***}$ | $0.041^{***}$ |
| | (2.168) | (4.353) | (3.187) |
| News-based uncertainty index | $-0.003^{***}$ | $-0.000$ | $-0.001^{***}$ |
| | $(-3.497)$ | $(-0.803)$ | $(-2.658)$ |
| U.S. dollar exch. rate (broad) | 0.004 | $0.044^{***}$ | $0.009^{**}$ |
| | (0.485) | (3.236) | (2.422) |
| Post-2008 $\times$ 10-year Treas. rate, pct | | | $-0.011$ |
| | | | $(-0.199)$ |
| Post-2008 | | | $0.498^{**}$ |
| | | | (2.193) |
| Dummy for crisis quarters | | | $-0.091^{*}$ |
| | | | $(-1.794)$ |
| Syndicate composition controls | $Yes$ | $Yes$ | $Yes$ |
| Lender type-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Lender country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Borrower industry-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Borrower country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Num. of observ. | 61507 | 18687 | 89655 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. within | 0.34 | 0.31 | 0.42 |
| RMSE | 0.31 | 0.18 | 0.32 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.

Note. The output shown is for regression models 2 and 4. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans made by all lenders to non-U.S. borrowers on an ultimate counterparty basis in the global market. Dependent variable is log(loan spread, pct) of loan $j$ in quarter $t$ made to borrower $b$ from country $b,c$ in industry $b,i$ by lender $l$ of type $l,i$ from country $l,c$. Syndicate composition controls, including the number of lenders in each loan, and fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

Table 5: Syndicate regressions: Loans made by U.S. lenders to non-U.S. borrowers

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Fed. funds rate, pct | −0.065** | | −0.049** |
| | (−2.325) | | (−2.357) |
| 10-year Treas. rate (z.c.), pct | 0.073 | −0.261*** | −0.013 |
| | (1.010) | (−3.730) | (−0.215) |
| Variance risk premium, pct sq. | 0.002 | 0.004 | −0.001 |
| | (0.796) | (0.806) | (−0.443) |
| European sovereign spread, pct | 0.030 | −0.121** | 0.007 |
| | (0.415) | (−2.384) | (0.195) |
| Expected inflation, pct | −0.061 | 0.113 | −0.008 |
| | (−1.212) | (0.820) | (−0.261) |
| Low-grade bond spread, pct | 0.189*** | 0.061 | 0.070*** |
| | (3.050) | (1.637) | (3.187) |
| News-based uncertainty index | −0.004*** | −0.002** | −0.001** |
| | (−3.119) | (−2.456) | (−2.443) |
| U.S. dollar exch. rate (broad) | −0.015 | 0.026 | 0.004 |
| | (−0.996) | (1.055) | (0.648) |
| Post-2008 × 10-year Treas. rate, pct | | | 0.044 |
| | | | (0.492) |
| Post-2008 | | | 0.394 |
| | | | (1.033) |
| Dummy for crisis quarters | | | −0.110 |
| | | | (−1.184) |
| Syndicate composition controls | $Yes$ | $Yes$ | $Yes$ |
| Lender type-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Lender country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Borrower industry-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Borrower country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Num. of observ. | 6588 | 2130 | 9568 |
| R-sq. w. | 0.40 | 0.41 | 0.45 |
| RMSE | 0.31 | 0.17 | 0.31 |
| RMSE | 0.30 | 0.18 | 0.32 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.
Note. The output shown is for regression models 2 and 4. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans made by U.S. lenders to non-U.S. borrowers on an ultimate counterparty basis in the global market. Dependent variable is log(loan spread, pct) of loan $j$ in quarter $t$ made to borrower $b$ from country $b, c$ in industry $b, i$ by lender $l$ of type $l, i$ from country $l, c$. Syndicate composition controls, including the number of lenders in each loan, and fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

Table 6: Syndicate regressions: Loans made by non-U.S. lenders to EME borrowers

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Fed. funds rate, pct | $-0.068^{**}$ | | $-0.079^{***}$ |
| | $(-2.567)$ | | $(-3.847)$ |
| 10-year Treas. rate, pct | 0.047 | $-0.210^{**}$ | $-0.001$ |
| | (1.017) | $(-2.588)$ | $(-0.034)$ |
| Variance risk premium, pct sq. | $0.005^{***}$ | $-0.004$ | 0.002 |
| | (3.097) | $(-0.692)$ | (1.375) |
| European sovereign spread, pct | 0.003 | $-0.129^{***}$ | $-0.024$ |
| | (0.087) | $(-3.475)$ | $(-1.036)$ |
| Expected inflation, pct | $-0.026$ | $0.201^{***}$ | 0.044 |
| | $(-0.574)$ | (3.986) | (1.593) |
| Low-grade bond spread, pct | $0.116^{***}$ | $0.129^{***}$ | $0.054^{***}$ |
| | (4.039) | (3.521) | (3.984) |
| News-based uncertainty index | $-0.004^{***}$ | $-0.001^{*}$ | $-0.001^{**}$ |
| | $(-4.174)$ | $(-1.985)$ | $(-2.452)$ |
| U.S. dollar exch. rate (EME) | 0.001 | $0.026^{**}$ | $-0.001$ |
| | (0.100) | (2.253) | $(-0.185)$ |
| Post-2008 × 10-year Treas. rate, pct | | | $-0.050$ |
| | | | $(-0.912)$ |
| Post-2008 | | | $0.534^{**}$ |
| | | | (2.281) |
| Dummy for crisis quarters | | | $-0.183^{***}$ |
| | | | $(-3.666)$ |
| Syndicate composition controls | *Yes* | *Yes* | *Yes* |
| Lender type-year fixed effects | *Yes* | *Yes* | *Yes* |
| Lender country-year fixed effects | *Yes* | *Yes* | *Yes* |
| Borrower industry-year fixed effects | *Yes* | *Yes* | *Yes* |
| Borrower country-year fixed effects | *Yes* | *Yes* | *Yes* |
| Num. of observ. | 45472 | 13736 | 66109 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. w. | 0.36 | 0.35 | 0.44 |
| RMSE | 0.30 | 0.18 | 0.32 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.

Note. The output shown is for regression models 2 and 4. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans originated in the global market to EME borrowers on an ultimate counterparty basis. Dependent variable is log(loan spread, pct) of loan $j$ in quarter $t$ made to borrower $b$ from country $b,c$ in industry $b,i$ by lender $l$ of type $l,i$ from country $l,c$. Syndicate composition controls, including the number of lenders in each loan, and fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

Table 7: Portfolio regressions: Portfolios of loans made by all lenders to non-U.S. borrowers

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Fed. funds rate, pct | $-0.028^{*}$ | | $-0.036^{***}$ |
| | $(-1.970)$ | | $(-3.419)$ |
| 10-year Treas. rate, pct | $0.040^{*}$ | $-0.219^{**}$ | 0.036 |
| | (1.690) | $(-2.560)$ | (1.555) |
| Variance risk premium, pct sq. | 0.002 | 0.002 | $0.002^{*}$ |
| | (1.590) | (0.623) | (1.975) |
| European sovereign spread, pct | $-0.015$ | $-0.049$ | $-0.037^{**}$ |
| | $(-0.637)$ | $(-0.963)$ | $(-2.292)$ |
| Expected inflation, pct | 0.005 | 0.088 | 0.021 |
| | (0.195) | (0.842) | (0.889) |
| Low-grade bond spread, pct | 0.019 | 0.032 | $0.037^{***}$ |
| | (0.611) | (1.329) | (3.799) |
| News-based uncertainty index | $-0.002^{**}$ | $-0.002^{**}$ | $-0.002^{***}$ |
| | $(-2.487)$ | $(-2.892)$ | $(-4.346)$ |
| U.S. dollar exch. rate (broad) | 0.002 | 0.016 | $-0.004$ |
| | (0.361) | (1.138) | $(-1.331)$ |
| Post-2008 $\times$ 10-year Treas. rate, pct | | | $-0.214^{***}$ |
| | | | $(-5.227)$ |
| Post-2008 | | | $1.327^{***}$ |
| | | | (8.806) |
| Dummy for crisis quarters | | | $-0.012$ |
| | | | $(-0.168)$ |
| Lender type-year fixed effect | *Yes* | *Yes* | *Yes* |
| Lender country-year fixed effect | *Yes* | *Yes* | *Yes* |
| Borrower country-year fixed effect | *Yes* | *Yes* | *Yes* |
| Num. of observ. | 14675 | 4791 | 21591 |
| Num. of lenders | 1690 | 675 | 1974 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. within | 0.15 | 0.12 | 0.32 |
| RMSE | 0.49 | 0.38 | 0.48 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.
Note. The output shown is for regression models 5 and 6. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans originated in the global market to non-U.S. borrowers on an ultimate counterparty basis. Dependent variable is log(average spread of loan portfolio, pct) of lender $l$ from country $l, c$ of type $l, i$ made to borrowers from country $b, c$ in quarter $t$. Fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

Table 8: Portfolio regressions: Portfolios of loans made by bank and nonbank financial lenders to non-U.S. borrowers

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Bank×Fed. funds rate, pct | $-0.029^{**}$ | | $-0.036^{***}$ |
| | $(-2.135)$ | | $(-3.705)$ |
| Nonbank×Fed. funds rate, pct | 0.000 | | $-0.016$ |
| | (0.009) | | $(-0.461)$ |
| Bank×10-year Treas. rate, pct | $0.044^{*}$ | $-0.224^{**}$ | $0.040^{*}$ |
| | (1.906) | $(-2.565)$ | (1.790) |
| Nonbank×10-year Treas. rate, pct | 0.008 | $-0.207^{**}$ | 0.010 |
| | (0.168) | $(-2.116)$ | (0.223) |
| Variance risk premium, pct sq. | 0.002 | 0.002 | $0.002^{**}$ |
| | (1.657) | (0.624) | (2.041) |
| European sovereign spread, pct | $-0.014$ | $-0.050$ | $-0.038^{**}$ |
| | $(-0.616)$ | $(-0.963)$ | $(-2.294)$ |
| Expected inflation, pct | 0.007 | 0.085 | 0.021 |
| | (0.246) | (0.808) | (0.909) |
| Low-grade bond spread, pct | 0.019 | 0.032 | $0.037^{***}$ |
| | (0.610) | (1.363) | (3.870) |
| News-based uncertainty index | $-0.002^{**}$ | $-0.002^{***}$ | $-0.002^{***}$ |
| | $(-2.544)$ | $(-2.942)$ | $(-4.356)$ |
| U.S. dollar exch. rate (broad) | 0.002 | 0.015 | $-0.004$ |
| | (0.383) | (1.069) | $(-1.355)$ |
| Bank×10-year Treas. rate, pct×Post-2008 | | | $-0.224^{***}$ |
| | | | $(-5.584)$ |
| Nonbank×10-year Treas. rate, pct×Post-2008 | | | $-0.147^{*}$ |
| | | | $(-1.959)$ |
| Post-2008 | | | $1.362^{***}$ |
| | | | (9.135) |
| Dummy for crisis quarters | | | $-0.011$ |
| | | | $(-0.153)$ |
| Lender type-year fixed effect | *Yes* | *Yes* | *Yes* |
| Lender country-year fixed effect | *Yes* | *Yes* | *Yes* |
| Borrower country-year fixed effect | *Yes* | *Yes* | *Yes* |
| Num. of observ. | 14098 | 4672 | 20844 |
| Num. of groups. | 1343 | 610 | 1563 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. within | 0.15 | 0.12 | 0.32 |
| RMSE | 0.49 | 0.38 | 0.48 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.

Note. The output shown is for regression models 5 and 6 that are modified to include lender type-specific regression coefficients. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans originated in the global market to non-U.S. borrowers on an ultimate counterparty basis. The number of observations is slightly lower than in Table 7 because nonfinancial lenders got excluded from the sample. Bank and Nonbank are dummy variables for banks and nonbank financial lenders respectively. Dependent variable is log(average spread of loan portfolio, pct) of lender $l$ from country $l, c$ of type $l, i$ made to borrowers from country $b, c$ in quarter $t$. Fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

## Appendix: Robustness checks

In this appendix, we cover some robustness checks in detail and mention some others in passing.

First, we note that while lenders and borrowers' risk management decisions are generally made at a parent level, some decisions to lend or borrow may be made on a local or immediate basis. To account for such a possibility, we reestimate regression models 2 to 6 on an immediate counterparty basis. In these regressions, we do not consolidate immediate borrowers and lenders under their respective parents. The results for syndicate regressions in Table A1 are very similar to those on an ultimate counterparty basis in Table 4 but the results for portfolio regressions in Table A2 are weaker than those in Table 7. The differences in statistical significance and explanatory power for portfolio regressions are not surprising because, again, risk management decisions tend to be made at a parent level. For example, on the supply side, a parent organization typically allocates capital towards risky assets and sets limits on credit risk of such assets.

Second, we note that our findings do not necessarily apply to syndicated term loans made to borrowers with investment-grade debt ratings. In fact, omitting loans to such borrowers from the sample gives statistically stronger results. In a way, we offer a falsification test where the cost-of-funds and returns-on-safer-assets channels associated with lower interest rates do not generally apply to loans made to investment-grade borrowers.

Third, we note that the results are stronger for most active lenders in the global syndicated market. The lenders that lend in the global syndicated loan market in every quarter tend to be large, sophisticated institutions that cater syndicated loans to a wide range of nonbank lenders, but their number in our sample is very modest.

Table A1: Syndicated regressions: Loans made by all lenders to non-U.S. borrowers on an immediate counterparty basis

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Fed. funds rate, pct | $-0.069^{***}$ | | $-0.082^{***}$ |
| | $(-2.966)$ | | $(-4.504)$ |
| 10-year Treas. rate (z.c.), pct | $0.075^{*}$ | $-0.110^{*}$ | 0.044 |
| | (1.941) | $(-1.927)$ | (1.122) |
| Variance risk premium, pct sq. | $0.003^{*}$ | $-0.003$ | 0.001 |
| | (1.810) | $(-0.739)$ | (0.506) |
| European sovereign spread, pct | 0.030 | $-0.092^{***}$ | 0.007 |
| | (1.086) | $(-3.367)$ | (0.292) |
| Expected inflation, pct | $-0.050$ | $0.156^{**}$ | $0.055^{*}$ |
| | $(-1.183)$ | (2.478) | (1.776) |
| Low-grade bond spread, pct | $0.117^{***}$ | $0.114^{***}$ | $0.049^{***}$ |
| | (2.768) | (5.005) | (3.352) |
| News-based uncertainty index | $-0.004^{***}$ | $-0.001^{*}$ | $-0.001^{***}$ |
| | $(-4.228)$ | $(-1.796)$ | $(-3.299)$ |
| U.S. dollar exch. rate (broad) | 0.008 | $0.027^{**}$ | $0.011^{***}$ |
| | (1.072) | (2.795) | (2.722) |
| Post-2008 × 10-year Treas. rate, pct | | | $-0.022$ |
| | | | $(-0.387)$ |
| Post-2008 | | | $0.680^{***}$ |
| | | | (2.767) |
| Dummy for crisis quarters | | | $-0.078$ |
| | | | $(-1.216)$ |
| Syndicate composition controls | *Yes* | *Yes* | *Yes* |
| Lender type-year fixed effects | *Yes* | *Yes* | *Yes* |
| Lender country-year fixed effects | *Yes* | *Yes* | *Yes* |
| Borrower industry-year fixed effects | *Yes* | *Yes* | *Yes* |
| Borrower country-year fixed effects | *Yes* | *Yes* | *Yes* |
| Num. of observ. | 61269 | 18542 | 89207 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. w. | 0.39 | 0.37 | 0.47 |
| RMSE | 0.26 | 0.15 | 0.27 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.

Note. The output shown is for regressions 2 and 4. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans originated in the global market to non-U.S. borrowers on an immediate basis. Dependent variable is log(loan spread, pct) of loan $j$ in quarter $t$ made to borrower $b$ from country $b,c$ in industry $i$ by lender $l$ of type $l,i$ from country $l,c$. Syndicate composition controls, including the number of lenders in each loan, and fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.

Table A2: Portfolio regressions: Portfolios of loans made by all lenders to non-U.S. borrowers on an immediate counterparty basis

| | (1) Pre-crisis | (2) Post-crisis | (3) Full |
|---|---|---|---|
| Fed. funds rate, pct | −0.014 | | −0.030*** |
| | (−1.185) | | (−3.073) |
| 10-year Treas. rate, pct | 0.065*** | −0.215** | 0.063*** |
| | (3.078) | (−2.275) | (2.871) |
| Variance risk premium, pct sq. | 0.002* | 0.003 | 0.002** |
| | (2.007) | (0.832) | (2.621) |
| European sovereign spread, pct | −0.016 | −0.031 | −0.033** |
| | (−0.818) | (−0.582) | (−2.132) |
| Expected inflation, pct | −0.034 | 0.122 | 0.029 |
| | (−1.144) | (1.018) | (1.074) |
| Low-grade bond spread, pct | 0.014 | 0.010 | 0.031*** |
| | (0.597) | (0.441) | (3.203) |
| News-based uncertainty index | −0.002** | −0.002** | −0.001*** |
| | (−2.456) | (−2.795) | (−2.975) |
| U.S. dollar exch. rate (broad) | 0.003 | 0.026 | −0.003 |
| | (0.593) | (1.595) | (−0.964) |
| Post-2008 × 10-year Treas. rate, pct | | | −0.237*** |
| | | | (−5.658) |
| Post-2008 | | | 1.514*** |
| | | | (10.139) |
| Dummy for crisis quarters | | | 0.027 |
| | | | (0.347) |
| Lender type-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Lender country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Borrower country-year fixed effects | $Yes$ | $Yes$ | $Yes$ |
| Num. of observ. | 25473 | 6947 | 35532 |
| Num. of groups. | 4373 | 1559 | 5155 |
| Num. of clusters | 50 | 18 | 76 |
| R-sq. w. | 0.10 | 0.11 | 0.23 |
| RMSE | 0.52 | 0.37 | 0.51 |

$t$ statistics in parentheses. $^{*}$ $p < .1$, $^{**}$ $p < .05$, $^{***}$ $p < .01$.
Note. The output shown is for regressions 5 and 6. Based on U.S. dollar-denominated, U.S. dollar LIBOR-indexed term loans originated in the global market to non-U.S. borrowers on an immediate basis. Dependent variable is log(average spread of loan portfolio, pct) of lender $l$ from country $l,c$ of type $l,i$ made to borrowers from country $b,c$ in quarter $t$. Fixed effects not shown. Column (3) includes crisis quarters from 2007:Q3 to 2009:Q2. Errors clustered by quarters.