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Bank Competition and Information Production

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Abstract

We show that bank competition diminishes banks' incentives to produce information about prospective borrowers. We exploit the deregulation of U.S. interstate branching as a shock to competition and use borrowers' stock returns after loan announcements to measure bank information production. Positive loan announcement returns are reduced in states that deregulate interstate branching, especially for opaque and bank-dependent firms and smaller banks that rely on soft information. Existing (i.e., inside) banks reduce information production more than new (i.e., outside) banks after deregulation, suggesting that they do so to deter borrower poaching. Furthermore, the probability of a covenant violation increases following deregulation.

I. Introduction

The issue of increasing competition in the banking industry is controversial. On the one hand, competition can improve access to credit and spur economic growth (e.g., Jayaratne and Strahan (1996), Cetorelli and Strahan (2006)). On the other hand, it can induce banks to extend loans to ex ante riskier borrowers (Carlson, Correia, and Luck (2022), Gissler, Ramcharan, and Yu (2020)). This can happen because of a reduction in bank charter value, which makes lenders more likely to reach for yield (Keeley (1990), Bisetti, Karolyi, and Lewellen (2020)), but also because competition can dilute banks' incentive to produce information about prospective borrowers. While an extensive literature has analyzed the impact of competition on bank risk-taking incentives, empirical evidence on whether bank competition can affect the information content of bank loans is scant.

In this paper, we provide novel evidence on the effects of competition on bank information production. Following James (1987), we use the cumulative abnormal return on the borrowing firm's stock after the announcement of a bank loan to

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measure the informativeness of a lending decision. We find that the information content of new loans in the syndicated loan market has decreased over time: Loan announcement returns were around 1% in 1980s and 0.35% in 1990s and 2000s. This downward trend, however, could be due to many different factors, such as the increase in bank competition or the adoption of new lending technologies. Therefore, to obtain plausibly exogenous variation in bank competition, we exploit the staggered deregulation of U.S. interstate branching between 1993 and 2006 (Rice and Strahan (2010)). Using data from DealScan, we find that the positive loan announcement returns are eliminated for firms headquartered in states that deregulate. Importantly, the results hold when we compare loan announcement returns for the same firm facing the same industry-time shock: Doing so rules out that firms obtaining loans in deregulating states are systematically different from firms in non-deregulating states.

We then test whether the effect of competition varies in the cross-section of borrowers' information sensitivity. Theory predicts that informationally opaque borrowers benefit more from lending relationships (Diamond (1991), Rajan (1992)), as banks can overcome the high degree of asymmetric information better than public markets. In line with this hypothesis, we find that loan announcement returns are positive and significant only for small, young firms and those without a bond rating (i.e., bank-dependent borrowers). Moreover, the removal of branching restrictions decreases loan announcement returns only of bank-dependent borrowers or those that operate in sectors with few tangible assets. Taken together, these results reinforce the notion that competition reduces banks' ability to produce information about borrowers for which information asymmetries are more likely to be a problem.

There are two potential explanations of our results. On the one hand, when competition increases, borrowers can take the loan offers they receive from existing (inside) banks to new (outside) lenders to bargain for better terms. In anticipation of this problem, inside banks optimally reduce their screening effort not to lose their best customers and avoid outside lenders free riding on their costly information production (Petriconi (2021)). On the other hand, new lending by outside banks, which potentially do not have the inside information to conduct a proper screening, could also lead to a decrease in the loan abnormal returns.[1]

We provide additional evidence to distinguish between the two explanations. First, we include lender characteristics and lender × post-deregulation fixed effect, thus controlling for lead arrangers' average screening ability and whether it changes after deregulation. We find that the effect of deregulation on loan announcement returns is unchanged. Second, we divide the sample between loan announcements from inside and outside lenders and find that the decrease in loan abnormal returns is primarily concentrated in the subsample of loans made by inside lenders. The evidence is therefore consistent with the hypothesis that inside

[1]One may question whether branching deregulation is a relevant shock to competition in the syndicated loan market because branching restrictions did not legally prohibit firms to obtain loans from out-of-state banks. In this respect, we find that interstate deregulation affected the composition of lead arrangers in syndicated loans over time: The share of deals syndicated by in-state lead arrangers decreased from 29% in 1993 to 9.6% in 2006.

banks reduce screening in response to an increase in competition to deter borrower poaching.

We also show that the reduction in information production due to the increase in competition has negative real effects on ex post loan performance. To do so, we test whether deregulation affected the probability of a covenant violation using the measure proposed by Demerjian and Owens (2016), a nonparametric assessment of the distance from the covenant threshold. Consistent with the hypothesis that loans made in a competitive environment are more likely to underperform, we find that the probability of a covenant violation, especially a performance covenant (e.g., interest coverage ratio), is higher in deregulated states.

We explore a rich set of alternative specifications to verify the robustness of our results. First, in a spirit similar to that of James (1987), who showed a positive stock response to loan announcements, but not to other public debt offerings, we use bond announcements over the same period as a placebo test and find that deregulation does not affect the stock price response after of a bond issue. Second, to address the shortcomings of staggered difference-in-differences (DiD) (Goodman-Bacon (2021), Baker, Larcker, and Wang (2022)), we use the "stacked" DiD methodology as in Cengiz, Dube, Lindner, and Zipperer (2019). Finally, we compute cumulative abnormal returns using the Fama and French (2015) 5-factor model instead of the 3-factor model and consider different loan announcement windows (1, 7, and 10 days) to calculate cumulative abnormal returns.

This paper is related to several different strands of the literature. In a seminal paper, James (1987) shows that bank lending is "special" in reducing asymmetric information because borrowers experience a positive abnormal stock return following the announcement of a bank loan and not for other public debt offerings.[2] More recently, Schwert (2020) finds that, after accounting for their higher seniority, bank loans earn a substantial premium over bonds, indicating that they have some "special" quality for the borrowing firms. We contribute to this literature by confirming that the positive loan announcement returns, initially documented by James (1987), is lower, but still present, 20 years later and it crucially depends on the degree of banking competition.

Our results also speak to the broader debate on the costs and benefits of competition in financial markets. A large theoretical literature has examined the effect of competition on banks' franchise values and risk-taking (Keeley (1990), Boyd and De Nicoló (2005)). Related to our work, Carlson et al. (2022) use data from the national banking era to show that competition increased bank risk-taking and default rates, Bisetti et al. (2020) argue that deregulation squeezed banks' net interest margins and increased charge-offs, and Gissler et al. (2020) show that competition in the consumer credit market led to an expansion of credit to riskier borrowers. This result is also related to the credit cycle literature, which argues that bank lending standards are countercyclical (Dell'Ariccia and Marquez (2006),

[2]The latter result is also found in Eckbo (1986). Lummer and McConnell (1989) further expand on James (1987) by showing that announcements of bank loan renewals, as opposed to new loans, are responsible for the positive stock response. Billett, Flannery, and Garfinkel (1995) document that banks with higher ratings are associated with larger borrower stock returns. Ongena and Roscovan (2013) find that loan announcement returns are higher when the loan is made by foreign bank.

Gorton and Ordonez (2018), Asriyan, Laeven, and Martin (2022), Cao, Di Pietro, Kokas, and Minetti (2022), and Weitzner and Howes (2023)), with some stressing the impact of competition on screening intensity (Ruckes (2004), Petriconi (2021)). We provide empirical evidence that competition can have additional adverse effects on asymmetric information because it lowers banks' incentives to screen and monitor.

We contribute to the literature on the role of banks as monitors in the syndicated loan market. Due to the lack of publicly available data, existing studies mostly rely on indirect proxies for monitoring, such as syndicate structure and the lead arrangers' share retention (Sufi (2007), Focarelli, Pozzolo, and Casolaro (2008), and Ivashina (2009)) or loan covenants (Sufi (2009)).[3] In this paper, the abnormal return after a loan announcement measures the value of both ex post monitoring and ex ante screening of borrowers. We show that the value of such information production depends on market power, an aspect that was overlooked in previous studies.

The latter result is related to Sharpe (1990) and von Thadden (2004), who argue that the bank informational monopoly may have adverse consequences on the allocation and cost of capital. Ioannidou and Ongena (2010) confirm these predictions empirically.[4] Differently from these papers, we directly test the role of market power in explaining banks' information advantage, exploiting an exogenous shock to bank competition.

II. Data and Stylized Facts

A. Banking Deregulation: Institutional Details

The U.S. has a long history of imposing restrictions on banks' ability to expand geographically, dating back to colonial times (Kroszner and Strahan (2014)). The U.S. states collected fees from granting charters only to banks headquartered within-state and so they had no incentives to foster competition from out-of-state banks. Between 1970 and 1994, U.S. began to gradually lift these restrictions in a staggered manner (Kroszner and Strahan (1999), Jayaratne and Strahan (1996)).[5] By the end of 1994, all states allowed intrastate branching and most allowed interstate branching, at least in principle. The formalization of the process at the national level happened in 1994, when Congress approved the IBBEA, the federal act that permitted banks to operate across state lines without any formal authorization from the target state.

In practice, though, interstate branching was still restricted even after 1994. The IBBEA, in fact, allowed individual states to erect barriers to prevent the entry

[3]An important exception is Gustafson, Ivanov, and Meisenzahl (2021), who directly measure the presence and frequency of active bank monitoring using confidential data. They confirm that share retention is a good proxy for bank monitoring, especially for more opaque firms.

[4]Similarly, Hale and Santos (2009) and Schenone (2010) find that when a firm goes for a bond or stock IPO, banks' informational rents decrease. Saidi and Zaldokas (2021) show that patent disclosure induces firms to switch banks, indicating that private information in banking relationships drops with the release of public information.

[5]Before 1970, bank holding companies could expand within state borders by setting up different bank subsidiaries but had to operate them separately. This severely limited the ability of banks to grow. For example, this meant that deposits could not be integrated in a single network and each subsidiary had to meet its own capital requirements (Kroszner and Strahan (1999)).

from out-of-state banks. In particular, states were permitted to use any number of the following 4 restrictions: i) mandating age restrictions on in-state banks that could be purchased (with a limit of no more than 5 years), ii) limiting the amount of deposits any new interstate merged bank could have in-state (up to 30%), iii) not allowing de novo interstate branching, and iv) not allowing the purchase of individual branches without acquiring the entire bank. These barriers mattered: The market share of deposits from out-of-state banks was only 2.5% in 1994, compared to 46% in 2011 (Keil and Müller (2020)). It also affected the composition of lead arrangers over time: The average fraction of deals syndicated by in-state lead arrangers decreased from 29% in 1993 to 9.6% in 2006.

Rice and Strahan (2010) create a count index of these restrictions in each state and year (from 0 to 4), where all states are considered to be fully restricted before 1993. We reverse the definition so that an increase in the index implies greater competition. The relaxation of the interstate branching restrictions was staggered across states, and Figure 1 shows the histogram of the number of deregulation changes using the first deregulation year in each state. There are 51 deregulation episodes for 43 states (8 states never deregulate). Most states (37 out of 43) deregulate between 1996 and 1998, and only 5 other states, including Texas, deregulate in 2000 and 2001. Therefore, despite the staggered nature of the deregulation, most changes occurred in just 3 years (1996–1998).

B. Loan Announcements

Our primary data source for bank loan announcements is LPC DealScan from 1993 to 2006.[6] For our purposes, we consider the issue date of the loan (DEAL_ACTIVE_DATE) as its announcement date.[7] We start by restricting the sample to U.S. borrowers, excluding those in the financial, real estate, and insurance sector (SIC codes 6000-6700). After applying these filters, we are left with 59,323 unique loan tranches (FACILITYIT) that we merge with Compustat using the link file provided by Chava and Roberts (2008). We are able to match 44,393 loan tranches to 6,742 firms (GVKEY) in Compustat.[8] We then use the CRSP-Compustat link file provided by WRDS to link GVKEY and PERMNO.

We download daily stock return data from CRSP around each loan announcement date. We set an estimation window of 150 days (and require companies to have at least 120 trading days of stock returns to enter the estimation) and a gap of 30 days before the announcement. We then run the Fama and French (1992) 3-factor model during the estimation window:

$$\mathrm{ER}_{i,t} = \alpha_i + \beta_{m,i}\mathrm{ER}_{m,t} + \beta_{SMB,i}\mathrm{SMB}_t + \beta_{HML,i}\mathrm{HML}_t \tag{1}$$

where $\mathrm{ER}_{i,t}=\mathrm{R}_{i,t}-\mathrm{R}_t^f$ is the excess return of stock i over the risk-free rate, $\mathrm{ER}_{m,t}$ is the market excess return and then calculate the abnormal returns as:

[6]We exclude the global financial crisis of 2007 and 2008, as it may confound our estimates.

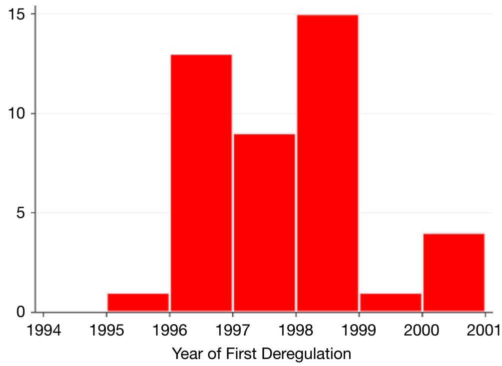
[7]Using a random sample of 250 announcements, we manually verified that around 90% of deal announcement dates on Dealscan coincide with the dates found on LexisNexis for the same deals.

[8]Private firms account for 91% of the unmatched loan tranches. Therefore, the coverage of publicly listed firms in Dealscan using the link file in Chava and Roberts (2008) is quite complete.

FIGURE 1

Number of Deregulation Changes (1993–2006)

Figure 1 contains the histogram of the number of changes in the Rice and Strahan (2010) deregulation index at the state level between 1993 and 2006.



$$\text{(2)} \qquad \mathrm{AR}_{i,t} = \mathrm{ER}_{i,t} - \left(\widehat{\alpha}_i + \widehat{\beta}_{m,i}\mathrm{ER}_{m,t} + \widehat{\beta}_{SMB,i}\mathrm{SMB}_t + \widehat{\beta}_{HML,i}\mathrm{HML}_t \right)$$

Finally, we compute the cumulative abnormal return for each announcement $\mathrm{CAR}_i = \sum_{\tau_1}^{\tau_2} \mathrm{AR}_{i,t}$ by summing the abnormal returns 5 days around the event date T $(\tau_1 = T - 1, \tau_2 = T + 3)$. For robustness, we also use the Fama and French (2015) 5-factor model, which adds operating profitability (robust minus weak) and investment (conservative minus aggressive) to the 3-factor model.

Additionally, we hand-collected loan announcements for the period 1980–1993, before DealScan data are available, using LexisNexis, an aggregator of news articles and feeds. We search the database for words such as "bank loan," "syndicate," "renewal," "expand," and "extend," as in Fields, Fraser, Berry, and Byers (2006). After cleaning and matching the data with CRSP, we obtained 723 loan announcement CARs between 1980 and 1992. Details on the data collection and cleaning are provided in Section A1 of the Supplementary Material.

C. Bond Announcements

We look for bonds issued by the firms with loan announcements in Dealscan and GVKEY-PERMNO links from CRSP-Compustat link file using the Mergent Fixed Income Securities Database (Mergent FISD). Mergent FISD contains information on the nature of the bond indenture including the amount issued, maturity, and the coupon. Similar to Dealscan, we consider the date when the bond was issued (OFFERING_DATE) as the bond announcement date. We are able to find 924 firms (using the issuer CUSIP as the firm identifier) that issue 4,346 bonds between 1993 and 2006. We then compute CARs for bond announcements in the same way as for loan announcements.

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D. Descriptive Statistics and CARs Over Time

The existing literature has shown that loan announcement returns were positive in the past but have declined in more recent decades. In the 1970–1984 period originally analyzed in James (1987), loan announcement CARs were 1.93%, but they have declined to 0.5% in 1990s and 0.1% in the early 2000s (Fields et al. (2006)).

First, we confirm that average loan announcement CARs are lower after the 1993 interstate deregulation. In Figure 2, we plot the average cumulative abnormal returns over a 20-day window around the announcement date separately for the 1980–1992 period, where we hand-collected data using Lexis Nexis, and the 1993–2006 period, where we collected loan announcements using Dealscan. On average, after obtaining a loan, a firm experiences a cumulative abnormal return of 1–1.5% in 1980–1992 but only about 0.4% in 1993–2006.

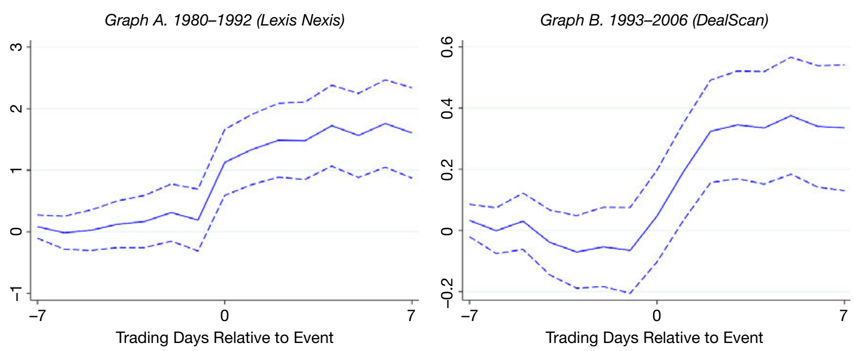
We estimate the difference in the CAR between the 2 periods more formally with a regression analysis in Table 1. The difference between the 2 periods is 0.74%; it is statistically significant at the 1% level and remains similar after controlling for the borrowers' industry and state fixed effects. The difference becomes larger (0.9%), even if less statistically significant, as we include firm fixed effects. The downward trend in average loan announcement CARs could be due to many different factors, such as the increase in bank competition or the adoption of different lending technologies. To identify the role of competition, we focus on the staggered interstate deregulation in the 1993–2006 period.

Table 2 provides the summary statistics for the variables used in the empirical estimation. Our final sample consists of 15,452 loan announcements from 3,015 unique firms between 1993 and 2006. Syndicated deals are large ($391 million on average) and have an average maturity of 3.75 years.[9] Because syndicated deals are

FIGURE 2

Loan Announcement Returns

Figure 2 plots the evolution of the average cumulative abnormal returns for a 20-day window around the loan announcement date.



[9]Deal maturity is a weighted average of the maturities of loan tranches within the same deal, with the weights equal to the share of the loan tranche out of the total deal value.

TABLE 1

CARs (1980–2006)

The dependent variable in Table 1 is the cumulative abnormal return around the loan announcement date, CAR(−1,3) from 1980 to 2006. Before 1993, the loan announcements come from LexisNexis and after 1993 from DealScan. Standard errors presented in parentheses are robust. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	1	2	3	4
I(YEAR <1993)	1.090***	1.162***	1.128***	1.247***
	(0.302)	(0.312)	(0.312)	(0.453)
I(YEAR ≥1993)	0.348***	0.347***	0.358***	0.344***
	(0.0522)	(0.0525)	(0.0529)	(0.0518)
Fixed effects				
Industry	No	Yes	Yes	-
State	No	No	Yes	-
Firm	No	No	No	Yes
Difference post-pre 1993	−0.742**	−0.815**	−0.770**	−0.902*
	(0.307)	(0.318)	(0.318)	(0.463)
No. of obs.	17,764	17,755	17,568	16,365
R^2	0.000	0.022	0.024	0.243

TABLE 2

Summary Statistics

Table 2 presents summary statistics for 15,452 loan announcements to 3015 Compustat firms by 90 lead arrangers (parent bank level) from 1993 to 2006 and 723 loan announcements from 1980 to 1992. RS_INDEX is the Rice and Strahan (2010) deregulation index. The definitions of firm controls are as follows: AGE is years since the first filing date with the SEC (first year in Compustat); $MKTCAP_{t-1}$ is market capitalization, stock price × common shares outstanding, PRCC_F× CSHO; BOOK_LEVERAGE is total debt/total assets, (DLC + DLTT)/AT; TANGIBILITY is net property, plant, and equipment/total assets, PPENT/AT; PROFITABILITY is operating income before depreciation/total assets, OIBDP/AT; CASH is cash holdings and short-term investments/total assets, CHE/AT; BOND_RATING is a dummy equal to 1 if the borrower has an S&P long-term issuer rating; SMALL_CAP is a dummy equal to 1 if the borrower is below the 20th percentile of the NYSE ME breakpoints (Fama and French (1992).

Variable	Obs.	Mean	Std. Dev.	25^{th}	50^{th}	75^{th}
CAR (1980–1992)	723	1.091	8.123	−3.444	0.277	4.406
CAR (1993–2006)	15,452	0.356	6.569	−2.836	0.030	3.178
CAR (bond)	4,346	0.477	4.409	−2.527	1.687	
RS_INDEX	15,452	1.516	1.523	0	1	3
Deal Controls						
DEAL_AMOUNT (USD mil)	15,452	391.15	868.56	50	150	400
DEAL_MATURITY (months)	15,452	45.40	29.51	24	40	60
Borrower Controls						
AGE (years)	15,452	21.814	16.719	7	15	36
$MKTCAP_{t-1}$ (USD mil)	15,452	5,061.1	1,9346.1	172.3	728.6	2,844.2
$BOOK\_LEVERAGE_{t-1}$	15,452	0.308	0.209	0.165	0.294	0.417
$TANGIBILITY_{t-1}$	15,452	0.352	0.240	0.155	0.295	0.528
$PROFITABILITY_{t-1}$	15,452	0.130	0.110	0.089	0.129	0.176
$CASH_{t-1}$	15,452	0.076	0.109	0.012	0.033	0.094
SMALL_CAP	15,398	0.606	0.489	0	1	1
BOND_RATING	15,452	0.661	0.473	0	1	1
Lead Arrangers Controls						
log(TOTAL_ASSETS)	7,340	12.422	1.133	11.712	12.502	13.373
TIER1_RATIO	7,340	8.201	1.344	7.500	8.150	8.500
CASH_ASSETS	7,340	0.112	0.049	0.078	0.110	0.141
DEPOSITS_ASSETS	7,340	0.576	0.128	0.538	0.589	0.661
LOANS_ASSETS	7,340	0.508	0.153	0.416	0.546	0.621

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large, the firms in our matched sample are on average older (22 years old) and have a larger market capitalization ($5 billion) and higher book leverage (30% of assets) than the average publicly listed firm in Compustat over the same period.[10] Still, we have a significant number of small-cap firms (SMALL_CAP = 1), that is, those with a total market capitalization below the 50th percentile of market capitalization in the NYSE as defined in the breakpoints from Fama and French (1992). These are 2,686 firms, representing 60% of the sample of loans. Furthermore, matching the sample of firms in Dealscan to Mergent FISD, we obtain 1,538 firms with a long-term S&P rating (BOND_RATING = 1), representing 66% of the deals. Not having access to the bond market or being a small-cap firm is a commonly used measure of bank-dependent borrowers (Chava and Purnanandam (2011), Schwert (2018)).

We obtain balance sheet information on the lead arrangers using the DealScan lender–Compustat link file provided by Schwert (2018). We restrict our attention to lead arrangers as these lenders are typically responsible for monitoring and due diligence on the loan (Focarelli et al. (2008), Ivashina (2009)). The link file matches DealScan lender names with Compustat GVKEY for all lenders with at least 50 loans or at least $10 billion in loan volume in the DealScan–Compustat sample. The link file also aggregates the lenders in DealScan at the parent bank holding company level, keeping track of mergers and acquisition. Most of the lenders in the DealScan sample are subsidiaries of large bank holding companies; hence, the lead arrangers (consolidated at the parent holding company) are large financial companies ($422 billion in total assets on average).

III. Empirical Strategy and Results

Banks' advantage in information acquisition rests on their ability to extract monopoly rents from their borrowers (Sharpe (1990), von Thadden (2004)). If banks face competition from other lenders, the information monopoly rents will be lower and inside banks will invest less in screening. Therefore, we expect the quality of the information produced for bank loans to be lower in areas where banks face more competition.

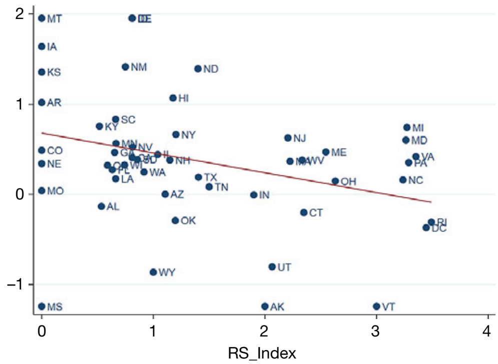
Figure 3 provides visual evidence on the relationship between average stock abnormal returns around loan announcement dates and the average deregulation index at the state level between 1993 and 2006. There is a negative relationship between the loan announcement return and the deregulation index: In more competitive states (i.e., a higher value of the index), the loan announcement return is lower. The simple regression line has a slope coefficient of -0.22 and is significant at the 5% level (t-stat = 2.41), with an R^2 of 0.10. Interestingly, the aggregate effect at the state level is remarkably close to the coefficient we estimate with the more granular data at the loan level below, where we exploit the different timing of the deregulation at the state level for the same firm. This suggests that the key cross-sectional differences in the level of the cumulative abnormal returns in our data are indeed driven by the deregulation episodes at the state level.

[10]The same figures for the universe of publicly listed firms on Compustat from 1993 to 2006 are 12 years old, $2 billion in market capitalization and 21% in book leverage.

FIGURE 3

Loan Announcement Returns and Deregulation (1993–2006)

Figure 3 provides a scatter plot of the state-average CAR after a loan announcement and the average of the RS_INDEX (Rice and Strahan (2010)) between 1993 and 2006.



More formally, we estimate the following specification:

$$\text{CAR}_{ifskt} = \beta_1 \text{RS\_INDEX}_{st} + \gamma' X_{ift-1} + \lambda_f + \lambda_{kt} + \varepsilon_{ifst} \quad (3)$$

where CAR_{ifskt} is the 5-day (−1,3) cumulative abnormal return for loan announcement i of firm f headquartered in state s operating in 2-digit industry k in year t. RS_INDEX_{st} is the deregulation index in state s in year t. X_{ift-1} is a vector of deal-specific and lagged firm-specific variables: loan maturity and size, book leverage, age, tangible to total assets ratio, profits, and cash holdings over assets. In our most saturated specification, we exploit the fact that the same firm obtains different syndicated loans over time and include a firm fixed effect, λ_f, together with 2-digit industry-year fixed effect, λ_{kt} (we include at least state and year fixed effects in all specifications, thus exploiting only variation across time for each deregulating state). Conditioning on these fixed effects yields a powerful test of our key hypothesis: The same firm, facing the same industry-time shock, experiences different abnormal returns depending on how the level of competition in the state varies over time. This rules out alternative risk-based explanations, such as that firms obtaining a loan are riskier in some unobservable way than others and hence carry a risk premium. Finally, we use wild-bootstrap standard errors based on state clusters following the methodology in Roodman, Nielsen, MacKinnon, and Webb (2019). Using bootstrap samples instead of clustering at the state level is appropriate given the small number of clusters (50 states) and large variability in the number of observations (loan announcements) within each cluster.

Table 3 reports our benchmark results. Across all specifications, the key coefficient of interest on the deregulation index is about −0.2, indicating that for each barrier to interstate branching that a state removes, the loan announcement returns decrease by 0.2%. Since on average states remove 2 barriers at a time when they deregulate (RS_INDEX = 2), the estimates imply that the deregulation fully eliminates the average loan announcement CAR (0.35%). Importantly, going from the combination of state and year fixed effects in column 1 to industry-time and firm

TABLE 3

CAR and Bank Competition

Table 3 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the loan announcement date, CAR(−1,3). RS_INDEX is the Rice and Strahan (2010) deregulation index. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	1	2	3	4	5	6
RS_INDEX	−0.197***	−0.188***	−0.217***	−0.223***	−0.218***	−0.223***
	(−2.751)	(−3.475)	(−3.443)	(−3.443)	(−3.443)	(−3.443)
MATURITY: 1–3 YEARS				0.224		0.227
				(1.038)		(1.057)
MATURITY: 3–5 YEARS				0.247		0.269
				(1.325)		(1.485)
MATURITY: > 5 YEARS				−0.096		−0.078
				(−0.376)		(−0.311)
AMOUNT: SECOND QUARTILE				−0.243		−0.199
				(−0.828)		(−0.691)
AMOUNT: THIRD QUARTILE				−0.255		−0.210
				(−0.946)		(−0.786)
AMOUNT: FOURTH QUARTILE				−0.599		−0.550
				(−1.477)		(−1.381)
BOOK_LEVERAGE					−0.224	−0.192
					(−0.459)	(−0.397)
log(1 + AGE)					−0.036	−0.026
					(−0.094)	(−0.067)
TANGIBILITY					0.098	0.077
					(0.090)	(0.070)
PROFITABILITY					−3.688***	−3.626***
					(−3.138)	(−3.152)
CASH					−1.053	−1.079
					(−0.832)	(−0.858)
Fixed effects						
State	Yes	Yes	-	-	-	-
Year	Yes	-	-	Yes	-	-
Industry-year	No	Yes	Yes	Yes	Yes	Yes
Firm	No	No	Yes	Yes	Yes	Yes
No. of obs.	15,452	15,452	15,452	15,452	15,452	15,452
R^2	0.004	0.056	0.289	0.290	0.291	0.291

fixed effects in column 3 has no effect on the coefficient on the deregulation index, suggesting that the deregulation is not correlated with firm unobservable characteristics or industry-specific shocks.

A. Maturity, Loan Size, and Leverage

We now consider some alternative explanations that may be driving our results. Competition could increase loan maturity or size, which would imply higher risk and leverage and hence lower CARs. To address this concern, in column 4 of Table 3, we control for quartile dummies of the distribution of loan maturities and amounts. None of these coefficients is statistically significant, and more importantly, the coefficient on the deregulation index is unchanged. Although not significant, the coefficients on maturity and size have the expected sign: Equity returns do not monotonically decrease with maturity but increase with short-term leverage and decrease with long-term leverage, consistent with the findings in Friewald, Nagler, and Wagner (2022). Larger loans, which increase leverage for a given firm size, instead decrease stock returns.

We control for firm characteristics, including book leverage, in column 5 of Table 3 and again find that the coefficient on the deregulation index is barely affected. The fact that the results are not affected by the inclusion of controls suggests that the treatment does not have strong heterogeneous impacts across subsamples (Baker et al. (2022)). The only firm characteristic that is negatively correlated with loan announcement returns is profitability.

B. Staggered DID and Other Robustness

To address the well-known bias of the 2-way fixed effects DiD estimator with staggered treatment (Goodman-Bacon (2021)) and in particular the "potentially problematic 2 × 2s" where early-treated units are used in the control group for late-treated units (Baker et al. (2022)), we adopt the "stacked regression" approach as in Cengiz et al. (2019). This method produces event-specific data sets that are "stacked" in event-relative time as opposed to calendar time and thus ensures that the set of controls associated with a treated unit are not treated during the time window. We show the results of the "stacked" DiD in Table A1 in the Supplementary Material and find that they are very similar to the baseline specification. In fact, despite the staggered nature of the interstate deregulation wave, most states deregulated between 1996 and 1998 (Figure 1), and therefore, the original data set is already quasi-"stacked."

To rule out that firm profits affect our results, we also reestimate abnormal stock returns using the 5-factor model (Fama and French (2015)), which adds operating profitability and investment factors to the 3-factor model. We find very similar results (see Table A2 in the Supplementary Material). Finally, we compute cumulative abnormal returns using different event windows of 3, 7, and 10 days around loan announcements instead of the baseline 5-day window (Table A3 in the Supplementary Material).

C. Placebo Test: Bond Announcements

A key result in James (1987) is that only announcements of bank loan agreements lead to a positive stock price response for the borrowing firm, whereas public debt offerings (or private debt placements to insurance companies) yield a nonsignificant or even negative stock reaction. He interprets these contrasting stock price responses as evidence that banks have access to inside information about the borrower that public markets or other intermediaries do not have, making bank loans "special."

In a similar spirit, we can use bond announcements by publicly listed firms in Dealscan over the same period as a placebo test. It may be that in the deregulated states other sources of funding become less informative too (i.e., there is nothing "special" about bank loans). We thus reestimate our specification in equation (3) for CARs computed around bond announcements instead of loan announcements. Results are presented in Table 4. We do not find any economically or statistically significant association between deregulation and bond abnormal returns,

TABLE 4

Bond CAR and Bank Competition: Placebo Test

Table 4 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the bond announcement date. RS_INDEX is the Rice and Strahan (2010) deregulation index. Deal and borrower controls are defined in Table 2. Wild bootstrap t-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	1	2	3	4	5	6
RS_INDEX	−0.021	−0.077	−0.064	−0.055	−0.068	−0.059
	(−0.249)	(−0.893)	(−0.607)	(−0.505)	(−0.600)	(−0.501)
Fixed effects						
State	Yes	Yes	–	–	–	–
Year	Yes	–	–	–	–	–
Industry-year	No	Yes	Yes	Yes	Yes	Yes
Firm	No	No	Yes	Yes	Yes	Yes
Borrower and deal controls	No	No	Yes	Yes	Yes	Yes
No. of obs.	4,346	4,346	4,346	4,346	4,346	4,346
R^2	0.017	0.118	0.318	0.321	0.320	0.324

suggesting that indeed deregulation affects only the market value of bank loans and not other sources of external funding.

Finally, a remaining concern with the estimates in Table 3 is that i) an unobserved factor, such as firm investment opportunities (i.e., higher credit demand), is affecting both the deregulation and the abnormal returns or ii) deregulation is affecting stock returns through alternative channels other than the increase in competition. We argue that this is unlikely to be the case. First of all, we note that if the deregulation improved firm investment opportunities, the abnormal return should be higher, not lower, as we find. Second, the interstate deregulation after 1994, as opposed to the intrastate deregulation in 1980s, has been shown not to affect the overall level of economic activity (Célérier and Matray (2019)).[11] In line with this and similar to findings in Keil and Müller (2020), we do not find evidence that deregulation increased credit supply in the syndicated loans, as neither rates decrease nor quantities increase following deregulation (Table A4 in the Supplementary Material).

D. Deregulation and the Entry of New Lenders

One interpretation of our results is that because of the increased competition from outside lenders, inside lenders reduce their screening effort to deter their existing customers from shopping around (Petriconi (2021)). However, the decline in loan abnormal returns is also consistent with the view that firms switch to outside lenders who do not have the inside information to conduct a proper screening because they are more distant from the borrowers (Granja, Leuz, and Rajan (2022)). In this section, we address these alternative explanations in two ways.

[11]This might be surprising in light of previous evidence that branching deregulation spurred economic growth Jayaratne and Strahan (1996). But those deregulation events (1977–1992) precede those analyzed here (1994–2006) and represent the first deregulation of regulatory constraints dating back to colonial times; thus, they are likely to have had the highest impact on real activity once they were removed. Also, while Kroszner and Strahan (1999) have shown that the strength of interest groups in the state can explain the first deregulation wave, the evidence for the interstate branching in the period 1993–2006 is not as clear (Rice and Strahan (2010)).

TABLE 5

CAR and Bank Competition: Lender Characteristics

Table 5 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the loan announcement date. RS_INDEX is the Rice and Strahan (2010) deregulation index. Deal and borrower controls are defined in Table 2. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	1	2	3	4	5	6
RS_INDEX	−0.172***	−0.205***	−0.210**	−0.190***	−0.197***	−0.278***
	(−3.148)	(−3.582)	(−2.511)	(−3.104)	(−3.951)	(−3.140)
log(TOTAL_ASSETS)	0.113	0.063	−0.081	0.007	−0.048	0.089
	(1.134)	(0.562)	(−0.457)	(0.027)	(−0.143)	(0.247)
TIER1_RATIO	−0.021	−0.033	−0.024	−0.071	−0.104	−0.140
	(−0.428)	(−0.747)	(−0.259)	(−0.855)	(−1.398)	(−0.843)
CASH_ASSETS	0.385	−0.910	−3.548	6.869***	5.104**	3.728
	(0.308)	(−0.850)	(−1.661)	(2.982)	(2.359)	(0.834)
DEPOSITS_ASSETS	−0.950	−1.298	−2.397*	−2.730	−3.232	−6.125**
	(−1.131)	(−1.225)	(−1.963)	(−1.368)	(−1.557)	(−2.274)
LOANS_ASSETS	0.886	0.787	1.500	3.484*	3.299*	6.766**
	(1.419)	(1.072)	(1.574)	(1.725)	(1.733)	(2.352)
Fixed effects						
State	Yes	Yes	–	Yes	Yes	–
Year	Yes	–	–	Yes	–	–
Industry-year	No	Yes	Yes	No	Yes	Yes
Firm	No	No	Yes	No	No	Yes
Lender-post	No	No	No	Yes	Yes	Yes
Borrower and deal controls	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	7,340	7,340	7,340	6,700	6,684	6,642
R^2	0.013	0.094	0.378	0.058	0.143	0.428

First, we include information on the identity of the lenders in the syndicate and show that lender characteristics are not driving our main results. In Table 5, we include lender (i.e., lead arranger) controls using the DealScan lender–Compustat link file from Schwert (2018).[12] Given the imperfect match with lead arrangers' balance sheet information, the sample is only half as large as in our baseline specification. There appears to be no systematic correlation between lender characteristics and cumulative abnormal returns on the loan, but, importantly, the effect of bank competition is virtually unchanged.

In columns 4–6 of Table 5, we include a lender × post fixed effect, that is, an indicator variable for each lender (or combinations of lenders if there is more than 1 lead arranger in the syndicate) both before and after the deregulation. In this case, we are controlling for the average screening ability of each syndicate and whether this changes together with the deregulation events. In all cases, we find that deregulation decreases loan announcement returns.

Second, we split our sample of loan announcements between those made by inside lenders (i.e., lead arrangers with which the borrowing firm had a preexisting syndicated loan) and those made by outside lenders. Nearly 40% of the loan announcements in our sample are made with lead arrangers with no preexisting relationships with the borrowers. We present the results in Table 6.

[12]Note that the unit of observation in all empirical specification is an individual loan (Packageid). Thus, for loans with more than 1 lead arranger (18% of the cases), the lender controls are simple averages of the lead arranger characteristics within each loan. We exclude all syndicates with more than 3 lead arrangers for simplicity (only 0.74% of loans have more than 2 lead arrangers).

TABLE 6

CAR and Bank Competition: Inside Versus Outside Lenders

Table 6 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the loan announcement date, CAR(−1,3). RS_INDEX is the Rice and Strahan (2010) deregulation index. Columns 1–3 restrict the sample to new loans from inside banks (i.e., lead arrangers with which the borrowing firm had a preexisting syndicated loan relationship), and columns 4–6 restrict the sample to new loans from outside banks (i.e., lead arrangers with no preexisting relationships with the borrowing firm). Deal and borrower controls are defined in Table 2. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	Inside Lenders			Outside Lenders		
	1	2	3	4	5	6
RS_INDEX	−0.280***	−0.391***	−0.316***	−0.092	−0.066	−0.238
	(−3.865)	(−4.230)	(−2.723)	(−0.772)	(−0.542)	(−0.617)
Firm, lender, and deal controls	Yes	Yes	Yes	Yes	Yes	Yes
Fixed effects						
Lender-post	Yes	Yes	Yes	Yes	Yes	Yes
State	Yes	Yes	-	Yes	Yes	-
Year	Yes	-	-	Yes	-	-
Industry-year	No	Yes	Yes	Yes	Yes	Yes
Firm	No	No	Yes	Yes	Yes	Yes
Borrower and deal controls	No	No	Yes	No	No	Yes
No. of obs.	4,127	3,971	3,052	3,777	3,621	1,908
R^2	0.077	0.212	0.470	0.089	0.201	0.672

We find that primarily the loans made by inside lenders experience a significant reduction in the abnormal returns after deregulation, whereas those made by outside lenders are not significantly affected by the increase in competition. Notably, when we include firm fixed effects, the coefficient on the deregulation index in the subsample of loans made by outside lenders is larger and similar to the baseline estimate, although still not statistically significant.

IV. Information Channel

Theory suggests that banks reduce agency costs associated with lending to opaque borrowers by screening and monitoring (Diamond (1991); Rajan (1992)). Thus, bank loans should be particularly "special" for informationally opaque borrowers (i.e., the loan announcement returns should be higher for opaque firms). Moreover, the loan announcement returns should decrease especially for informationally opaque firms after deregulation, since the incentives to screen opaque borrowers decrease as lenders face more competition. We find that both predictions are true in the data.

We explore several measures of firm opaqueness: the ratio of tangible to total assets of the sector in which the firm operates, the market cap of the firm, its age, and whether it has access to the bond market. While the first measure is the most directly related to information (i.e., firms with more tangible assets require less screening and have more collateral), the last 3 (i.e., small, young firms and those with no bond rating) are measures of bank dependence, which are also used as proxies for information opaqueness (Schwert (2018)).

First of all, in Figure 4, we document that the average cumulative abnormal return after obtaining a loan is positive and significant only for small-cap firms or those with no bond rating. The average CARs are about 0.6% for bank-dependent

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FIGURE 4

Loan Announcement Returns: Firm Heterogeneity

Figure 4 plots the evolution of the average cumulative abnormal returns for a 20-day window around the loan announcement date for bank-dependent firms. Graph A splits the sample into small-cap firms, defined as those with market capitalization below the NYSE breakpoints (blue line) and all other firms (red line). Graph B splits the sample into firms without a bond rating (blue line) and those with a bond rating (red line).

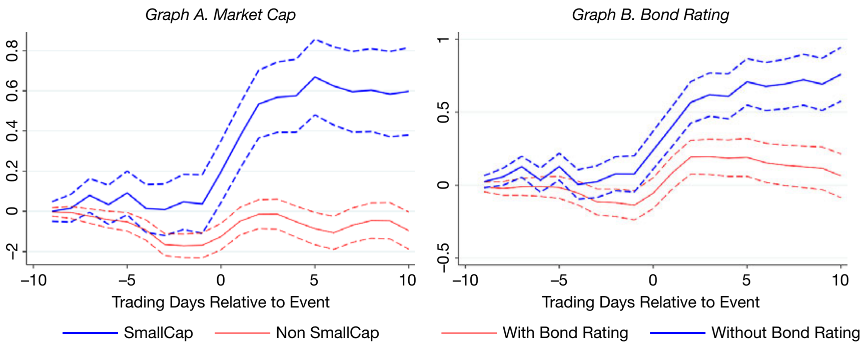

TABLE 7

Firm Opaqueness: Bank-Dependent Borrowers

Table 7 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the loan announcement date. RS_INDEX is the Rice and Strahan (2010) deregulation index. Small is a dummy equal to 1 for firms with below the median market capitalization in the NYSE as defined by the breakpoints in Fama and French (1992). Young is a dummy equal to 1 for firms with below the median age in the sample (15 years old). Bond rating is a dummy equal to 1 for firms with a long-term bond rating from S&P. Deal and borrower controls are defined in Table 2. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	Small		Young		Bond Rating	
	No	Yes	No	Yes	Yes	No
	1	2	3	4	5	6
RS_INDEX	−0.006	−0.332**	−0.123	−0.414**	−0.155*	−0.379*
	(−0.054)	(−2.411)	(−1.608)	(−2.509)	(−1.664)	(−1.878)
Fixed effects						
Industry-year	Yes	Yes	Yes	Yes	Yes	Yes
Firm	Yes	Yes	Yes	Yes	Yes	Yes
Borrower and deal controls	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	5,784	9,126	7,524	7,538	10,157	5,114
R^2	0.264	0.355	0.295	0.353	0.242	0.413

borrowers (50% higher than in the baseline), whereas they are close to 0 and nonsignificant for large firms or those with bond ratings. This suggests that bank loans are “special” only for bank-dependent firms.

Second, we test whether the effect of competition on the abnormal returns varies across firm characteristics. The results for bank-dependent borrowers are shown in Table 7. In general, we find that only bank-dependent borrowers experience a decrease in the abnormal return associated with a loan announcement after an increase in competition. The effect for small-cap firms is 50% larger than the baseline estimate, and it is virtually 0 and nonsignificant for mid- to-large-cap firms. Similarly, young firms (i.e., those with age below the sample median (15 years old)) experience a decline in abnormal returns that is twice as large as the baseline

TABLE 8

Firm Opaqueness: Tangibility

The dependent variable in Table 8 is the cumulative abnormal return around the loan announcement date. RS_INDEX is the Rice and Strahan (2010) deregulation index. TANG_RATIO is the ratio of tangible to total assets (PPENT/AT) at the 2-digit SIC code in which the firm is operating. Columns 3 and 4 restrict the sample to industries with, respectively, below and above the median of the tangibility ratio across industries. Deal and borrower controls are defined in Table 2. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

			TANG_RATIO	
			Low	High
	1	2	3	4
RS_INDEX	−0.394***	−0.426**	−0.310**	−0.183*
	(−3.422)	(−2.372)	(−2.331)	(−1.647)
RS_INDEX × TANG_RATIO	0.582**	0.548		
	(2.283)	(1.234)		
Fixed effects				
State	Yes	–	–	–
Industry-year	Yes	Yes	Yes	Yes
Firm	No	Yes	Yes	Yes
Borrower and deal controls	Yes	Yes	Yes	Yes
Borrower and deal controls × TANG_RATIO	Yes	Yes	–	–
Test of coefficients				
RS_INDEX + RS_INDEX × TANG_RATIO = 0.34	−0.196***	−0.243***		
	(0.055)	(0.058)		
No. of obs.	15,452	15,452	7,625	7,658
R^2	0.060	0.293	0.317	0.285

estimates, whereas older firms do not have any significant change. Finally, firms that do not have a bond rating experience a decline in the loan announcement abnormal returns after deregulation that is twice as large as those with a bond rating. Given that bank-dependent firms are those with higher CARs to start with (Figure 4), the results confirm that deregulation eliminates the positive stock response to loan announcements.

The results using tangibility as a measure of firm opaqueness are presented in Table 8. In column 1, we interact the deregulation index for the ratio of tangible to total assets at the 2-digit sector level in which the firm operates. The interaction term is positive and significant, indicating that in sectors with more tangible assets, where less information production is necessary, the effect of competition on loan returns is diminished. For example, the effect of deregulation is twice as large (−0.400) for a firm operating in a hypothetical industry with no tangible assets compared to one operating in an industry at the mean tangibility ratio (−0.196). The interaction term is not significant, but the point estimate remains similar when we include firm fixed effects in column 2. We also find that the effect of deregulation is 50% larger in sectors with below the median tangibility sector (column 3) compared to sectors above the median (column 4), for which the estimated effect is similar to the baseline estimate.

Finally, we expect competition to drive down the abnormal return of loans made by "relationship lenders," that is, banks whose business model relies on building lending relationships and invest more in "soft" information acquisition about their borrowers. Given that empirical measures of whether a bank is a relationship lender and measures of bank screening intensity are not easily observed

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TABLE 9

CAR and Bank Competition: Large Versus Small Banks

Table 9 provides estimates for equation (3). The dependent variable is the cumulative abnormal return around the loan announcement date. RS_INDEX is the Rice and Strahan (2010) deregulation index. LARGE_BANK is a dummy equal to 1 for lead arrangers with above median total assets in each year. Deal and borrower controls are defined in Table 2. Wild bootstrap *t*-statistics are presented in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

				LARGE_BANK	
				No	Yes
	1	2	3	4	5
RS_INDEX	−0.245***	−0.276***	−0.280***	−0.245**	−0.068
	(−3.651)	(−3.901)	(−3.159)	(−2.244)	(−0.493)
RS_INDEX × LARGE_BANK	0.150*	0.148*	0.148*		
	(1.764)	(1.682)	(1.744)		
LARGE_BANK	0.108	0.124	0.341		
	(0.473)	(0.520)	(1.057)		
Fixed effects					
State	Yes	Yes	-	-	-
Year	Yes	-	-	-	-
Industry-year	No	Yes	Yes	Yes	Yes
Firm	No	No	Yes	Yes	Yes
Borrower and deal controls	Yes	Yes	Yes	Yes	Yes
No. of obs.	7,336	7,336	7,335	3,074	3,163
R^2	0.013	0.094	0.378	0.484	0.457

(Gustafson, Ivanov, and Meisenzahl (2021)), we use bank size as a proxy for soft information acquisition (Berger, Miller, Petersen, Rajan, and Stein (2005)). The results are presented in Table 9.

As expected, we find that the effect of deregulation on abnormal returns after a loan announcement varies with bank size: In column 1, the baseline effect for a bank with below the median total assets in each year is negative and significant, and it is reduced by 60% (0.15/0.245) for large banks with above the median assets. The estimated effect remains unchanged as we saturate the regression with firm and industry-year fixed effects (columns 2 and 3). More explicitly, if we use a sample split for banks below and above the median assets, we find a negative effect of competition on the loans made by banks with total assets below the median, not for those above the median (columns 4 and 5). The results thus indicate that presumably only banks that are better able to produce soft information are affected by changes in competition.

So far, we have shown that competition decreases the information content of bank loans by decreasing bank incentives to invest in information production. An implication of our hypothesis is that the quality of loans originated after deregulation is lower (i.e., ex post defaults are higher in deregulated states). We can test whether this is the case in the data. Unfortunately, syndicated loans from DealScan do not allow to track the performance of the loan over time. To this end, we analyze another aspect of loan performance: the probability of violating a covenant at the loan inception date. The higher the probability of a covenant violation for a loan originated after deregulation, the lower the quality of the loan or the firm receiving it.[13]

[13]An alternative interpretation is that the lenders are requesting stricter conditions on financial covenants (i.e., a supply side restrictions). This interpretation is unlikely in our setting, since if anything we expect supply conditions to have eased, rather than tightened, as a consequence of deregulation.

TABLE 10

Covenant Violations and Bank Competition

The dependent variable in Table 10 is the probability of a covenant violation for a syndicated loan from Demerjian and Owens (2016). In column 1, the dependent variable is the aggregate probability of any covenant violation, while columns 2 and 3 use the probability of a performance and capital covenant violation, respectively. RS Index is the Rice and Strahan (2010) deregulation index. Wild bootstrap *t*-statistics in parentheses. *, **, and *** denote significance at 10%, 5%, and 1%, respectively.

	Any Covenant	Performance Covenant	Capital Covenant
	1	2	3
RS_INDEX	0.014**	0.014**	0.001
	(2.159)	(2.362)	(0.143)
Fixed effects			
Industry-year	Yes	Yes	Yes
Firm	Yes	Yes	Yes
No. of obs.	10,199	10,199	10,199
R^2	0.584	0.601	0.531

We use the probability of a covenant violation measured as a nonparametric distance from the covenant threshold from Demerjian and Owens (2016), who build a generalization of the loan strictness measure in Murfin (2012). The results are presented in Table 10.

We find that, for a state that fully opens up to competition, the probability of a covenant violation goes up by 6 percentage points (4×0.015), a 15% increase compared to the mean. Moreover, when we separate the probability of violation between performance and capital covenants, we find that only the former are responsible for the increase after deregulation. This is in line with the hypothesis that loan quality decreases following deregulation. Violating performance covenants, which are related to minimum interest coverage ratios, indicates that the firm is more likely to default because it cannot service its debt obligations. Capital covenants, on the other hand, are mostly related to short-term liquidity (e.g., quick ratio) or overall leverage.

V. Conclusion

A large literature argues that competition can have both positive and negative effects on bank risk-taking and performance. We argue that an important aspect that has been overlooked in previous studies is the fact that market power is an important determinant of lenders' incentives to produce information. In fact, competition decreases the ability to extract future rents from the borrowers and may thus decrease lenders' incentives to screen and monitor. Empirically, we observe that exogenous increases in local bank competition reduce the positive abnormal return associated with bank loan announcements. The decline in loan announcement returns occurs only for informationally opaque borrowers or banks that rely more on soft information. Moreover, we find that the probability of violating covenants increases in areas that open up to competition.

Our findings are informative regarding the cost and benefits of competition in financial markets. They suggest that bank competition, by limiting the ability of

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lenders to appropriate gains from informed lending, may deteriorate credit quality and exacerbate asymmetric information problems.

Supplementary Material

To view supplementary material for this article, please visit http://doi.org/10.1017/S0022109024000152.

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