Interest Rate Caps, Corporate Lending, and Bank Market Power: Evidence From Bangladesh*

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Abstract

How does market power in corporate bank lending influence the effects of interest rate cap policies on credit allocation in a low-income country? We study this question using credit registry data from Bangladesh, where the central bank capped rates on corporate loans between 2009 and 2011. Using difference-in-differences designs with variation in preregulation branch-level interest rates, we find that a one-percentage-point cap-induced drop in rates increased lending by 36%. This increase in lending did not drive up delinquency rates. Our results point to underprovision of credit due to banks' ex ante market power before establishing relationships with borrowers.

Keywords: corporate banking, interest rate caps, bank market power, regulation, emerging markets finance

JEL classifications: D43, G21, G28, L51, O16

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1 Introduction

A well-functioning financial market promotes economic growth by efficiently allocating credit to firms that need it (Rajan and Zingales 1998, Levine 2005, Midrigan and Xu 2014). However, firms in developing countries often face high external financing costs (Demirgüç-Kunt et al. 2004, Beck et al. 2005, Cavalcanti et al. 2024). To what extent are these elevated corporate lending rates driven by the inherent costs and risks banks face in offering loans, and to what extent do market frictions, such as banks' market power, play a role? Can policy interventions, such as interest rate caps, effectively lower lending rates and expand firms' access to credit?

This paper empirically assesses these questions using an interest rate cap policy as a natural experiment. Imposing a ceiling on corporate lending rates is one of the most widely used tools by regulators in emerging markets to alleviate firms' financing constraints and spur corporate investment (Maimbo and Henriques Gallegos 2014, Ferrari et al. 2018). We use credit registry microdata and a 2009 policy reform in Bangladesh, where the central bank imposed a maximum limit for interest rates on loans made to large and medium-scale enterprises at 13% (from the precap period average of 14.5%), under the stated objective of boosting industry investment in the aftermath of the Global Financial Crisis (Unnayan Onneshan 2011). We study how this interest rate cap regime affects credit provision and market outcomes during the regulation period.

To guide our empirical analysis, we develop a simple two-period model to obtain testable predictions for the effects of interest rate caps on equilibrium credit provision when banks face imperfect competition. In the first period, banks face a continuum of entrepreneurs looking to finance their projects. Banks set interest rates depending on the cost of funds, the risk of default, the expected value of future relationships, and the degree of competition that we call *ex ante market power*. When a borrower successfully repays their loan, banks can lock in the borrower and extract a profit surplus from the relationship in the second period, depending on the strength of *ex post market power*, as in Petersen and Rajan (1995). We show that the impact of an interest rate cap on equilibrium credit provision depends on both ex ante and ex post forms of market power. When ex ante market power is large, the interest rate cap tends to increase equilibrium credit provision as long as the cap is above the bank's break-even rate. When ex ante market power is low, the interest rate cap is more likely to induce credit rationing, even if ex post market power is high.

To resolve this theoretical ambiguity, we draw on Bangladesh's 2009 interest rate cap policy. This policy induced sudden changes in interest rates for corporate loans. We show how this policy change induced an increase in credit provision.

A key challenge in identifying the causal effects of this type of regulation is that rate caps are often endogenously enacted when credit market conditions and the macroeconomic outlook are less favorable. Hence, simply relying on time-series variation comparing outcomes across prereform and postreform periods would be inconclusive. To address this issue, we use preregulation interest rates at the bank-branch level as a source of plausibly exogenous cross-sectional variation. In particular, we adopt a difference-in-differences (DiD) research design in which we compare branches within the same parent bank that were more vs. less "exposed" to the 13% rate cap, depending on average rates charged on short-term corporate loans in the lead-up to the reform.

We show that bank branches charging high interest rates relative to the cap threshold prior to the regulation held interest rates stable until the policy was implemented, then suddenly lowered their interest rates after the policy's introduction. On the other hand, for inframarginal bank branches whose interest rates were already lower than the cap prior to its implementation, interest rates remain stable throughout the cap regime. Importantly, these patterns exist even after we include bank-quarter fixed effects, thereby controlling for any time-varying bank-level cost and demand shocks. We use these sudden, differential changes in branch-level interest rates to identify the causal effects of interest rate caps on equilibrium credit provision.

Our main finding is that bank branches more exposed to the interest rate cap experienced a significant *increase* in credit provision compared to less-exposed branches. This result is consistent with the interpretation that banks possess substantial ex ante market power. The expansion in credit provision occurred along both the extensive margin (an increase in the number of loans) and the intensive margin (an increase in the average loan size). The effect is economically large: a one-percentage-point cap-induced reduction in interest rates led to an increase of 31 log points (approximately 36%) in the total outstanding loan amount. Both the extensive and intensive margins contribute to this increase, while the latter makes a slightly larger contribution to the overall increase in lending.

The implied elasticity of credit demand with respect to the real interest rate is approximately 1.7. This value is higher than existing estimates of credit demand for small or microfinance loans in low-income countries based on experimental/quasi-experimental methods (Karlan and Zinman 2008, Dehejia et al. 2012). This is consistent with the view that larger firms face more elastic demand due to having access to alternative sources of financing, such as bond markets. At the same time, the elasticity is low enough that banks can still exercise substantial ex ante market power.

Our event-study analysis shows no evidence of pretrends in branch-level credit provision prior to the introduction of the cap and reveals a sharp change in lending behavior immediately afterward, supporting the parallel-trends assumption. To further address concerns about potential violations of this assumption, we conduct a placebo test using loans to individuals. Since the cap applied exclusively to corporate loans, it should not affect credit outcomes in the individual loan segment. Reassuringly, we find no impact on interest rates or credit provision for individual loans, giving additional support to our identification strategy. We also assess the robustness of our results using the test proposed by Rambachan and Roth (2023), which explicitly allows for potential violations of the parallel-trends assumption. Even under conservative assumptions about such violations, our results remain statistically significant.

We also assess how this increase in credit provision varies across branches with different precap characteristics. The effects are similar across branches operating in markets with varying levels of concentration in terms of lending amounts, as measured by the Herfindahl-Hirschman Index (HHI), indicating that credit rationing did not occur even in relatively competitive local markets. We find stronger effects in more populous regions of Bangladesh, consistent with higher credit demand elasticities in urban areas. The effects are also comparable across branches with differing levels of risk, as proxied by precap delinquency rates and the share of loans secured by physical collateral, as well as across branches with different initial deposit rates. We also observe statistically similar effects across branches belonging to banks with different levels of financial solvency, as measured by bank-level leverage ratios and delinquency rates.

Our empirical findings contrast sharply with other existing studies on lending rate caps focusing on high- or upper-middle-income countries and those imposed on loans to consumers or small firms. By implementing similar DiD designs, existing studies typically find a *decrease* in credit provision, particularly for riskier borrowers.[1] In contrast, we find an *increase* in credit provision in response to a rate cap. We argue that this contrasting conclusion stems from two key differences in our policy environment relative to the literature. One, we study a low-income country, Bangladesh, where the capital market is less mature than those in the existing studies.[2] Two, we focus on a rate cap applied to corporate loans for larger firms, for which a limited number of banks may have been able to extend loans and thus exercise a significant degree of market power.

We also investigate the branch-level effects of the cap on three other sets of credit market outcomes. First, we study the impacts on the risk composition of the borrower pool, as proxied

[1]For example, Benmelech and Moskowitz (2010), Rigbi (2013), Fekrazad (2020), and Cherry (2024) study caps on consumer and payday loans in the United States; Madeira (2019) and Cuesta and Sepulveda (2021) study consumer-loan rate caps in Chile; Alessie et al. (2005) does the same in Italy; and Burga et al. (2024) study caps on loans to small firms in Peru. With the exception of Alessie et al. (2005) and Fekrazad (2020), these studies uncover a *decrease* in credit provision.

[2]Bangladesh was classified as a low-income country until 2015, when it moved up to a lower-middle-income country (World Bank 2015). In 2008, GDP per capita in Bangladesh was nearly one-sixth that of Peru, the lowest-income country studied in this literature in which there was a negative effect of an interest rate cap on credit provision (Burga et al. 2024).

by ex post delinquency rates, the proportion of secured loans, and the composition of borrower sectors, which vary in their delinquency rates. We observe no effects on any of these outcomes, supporting our interpretation that the cap did not induce credit rationing to riskier borrowers. Again, this evidence is in sharp contrast with previous studies in the context of loans to consumers or small firms in middle- or high-income countries, which find that the cap leads to a shift of credit away from riskier borrowers (e.g., Cuesta and Sepulveda 2021; Burga et al. 2024; Cherry 2024), and bolsters evidence in favor of ex post market power within the riskier segment of borrowers.

Second, we examine the effects of the cap on the deposit market. If banks possess market power not only in lending but also in deposit-taking, as observed in the literature on the deposit franchise in the United States (Drechsler et al. 2017), they may respond to the lending rate cap by adjusting deposit rates. However, we find no statistically significant changes in deposit rates or amounts, either at the overall branch level or specifically among corporate depositors. Combined with the absence of significant changes in risks mentioned above, our results suggest that bank branches experienced a reduction in profit margins due to the interest rate cap.

Third, we investigate whether the cap induces branches to reallocate credit across different industries of borrowers. We find no statistically significant effects of the cap on the share of credit provided to different sectors. This finding implies that the cap expanded credit access equally across all sectors. This is an important finding given the outsize importance of financing for import-export firms in Bangladesh and other lower- and middle-income countries, as documented by Paravisini et al. (2015) and Paravisini et al. (2023).

Given that the introduction of the interest rate cap facilitated an increase in credit provision, our results are consistent with the presence of ex ante market power by banks. To further support our interpretation that banks wield ex ante market power, we provide two additional analyses. One, using the same interest rate cap policy in Bangladesh, we study how the reduction of interest rates by bank branches belonging to a close-competitor bank—which we define as a bank with similar balance-sheet size and sectoral targeting of loans—and situated in the same local market affects credit provision. If bank branches are in direct competition for borrowers, a reduction in interest rates by competing bank branches affects a branch's own credit demand. We estimate that the own-branch effects of a cap-induced interest rate reduction by competing bank branches are close to zero, with tight standard errors. This evidence is consistent with the interpretation that bank branches are under local monopoly, and that changes in interest rates by competitors do not affect their overall credit provision.

Second, we examine the impacts of the opening of a new branch on local competitors' interest rates and credit provision. Similarly to the previous idea, if bank branches face fierce competition for borrowers, the entry of a bank branch can attract borrowers away from incumbent branches, affecting equilibrium credit provision and interest rates. We fail to find statistically significant effects of the entry of competing bank branches on incumbent branches' credit market outcomes. This result holds even though we implement the estimator proposed by de Chaisemartin and D'Haultfœuille (2020), which identifies treatment effects using comparisons only between branches treated at different times, ruling out never-treated branches. Our evidence is again consistent with an interpretation that bank branches are under local monopoly, and that the entry of competitors does not affect their equilibrium credit provision or interest rates.

Our results offer clear evidence that banks in Bangladesh possess substantial ex ante market power, thus resulting in underprovision of credit prior to the cap. However, we caution against broad policy conclusions about interest rate cap policies. Our DiD approach identifies branch-level effects and does not capture certain general equilibrium effects, such as changes in country-level deposit rates, firm investment spillovers, or cross-market spillovers. In addition, our analysis focuses solely on short-run effects of a rate cap on credit provision; it does not address potential interactions between interest rate caps and monetary policy, or longer-term impacts on bank behavior such as branch network expansion. Nevertheless, our results underscore the need for policymakers to weigh ex ante market-power distortions when determining the desirability of credit market regulation. Interest rate caps may be one possible, though not exclusive, policy tool to address this type of distortion in corporate lending.

We contribute to various literatures on banking and finance, particularly work studying the financial sector in low-income countries, by providing evidence for banks' ex ante market power using an interest rate cap policy as a natural experiment.

Our work contributes to the literature investigating the causal impacts of interest rate caps by providing evidence from a low-income country on large corporate loans. As mentioned above, existing work on the causal effects of interest rate caps is largely limited to high- and middle-income countries, as well as caps applied to loans to consumers or small firms. In those contexts, studies often report a *decrease* in credit provision (Benmelech and Moskowitz 2010, Rigbi 2013, Madeira 2019, Cuesta and Sepulveda 2021, Burga et al. 2024, Cherry 2024). Our work demonstrates that, in a low-income country and for large-scale corporate lending, the opposite may be true, suggesting that banks may have substantial ex ante market power.[3]

Our work also contributes to the literature on barriers to financial access in low-income countries by highlighting banks' market power as a key distortion. Existing work has established evidence that firms in developing countries face severe credit constraints (De Mel et al. 2008, Hsieh and Klenow 2009, Kaboski and Townsend 2011, Banerjee and Duflo 2014, Bau and Matray 2023), and the expansion of bank branches alleviates these constraints and helps expand local economic and firm activity (Burgess and Pande 2005, Ji et al. 2023, Fonseca and Matray 2024). However, little is known about how efficiently banks provide credit once they enter the market. Our evidence suggests that market power in the banking sector is a key force that can lead to the underprovision of credit.

Finally, our paper contributes to the growing empirical literature documenting banks' ex ante market power. While ex post market power in relationship lending is well-established (e.g.,

[3]Interest rate caps on loans originated by microfinance institutions (MFIs) are also a common type of cap policy in developing countries. Quirk (2023) studies a cap that broadly targeted all microfinance loans in Zambia and finds large negative effects on credit provision. Heng et al. (2021) study an MFI loan cap in Cambodia and argue that MFIs increase their origination fees to make up for lost profits from the cap, while the number of loans declines.

Petersen and Rajan 1995, McMillan and Woodruff 1999, Beck et al. 2018), direct evidence on ex ante market power remains relatively scarce. Some studies infer it from cross-sectional differences in bank profit margins across countries, regions, or firms (Demirgüç-Kunt et al. 2004, Schwert 2020, Cavalcanti et al. 2024), while others exploit the heterogeneous pass-through of monetary policy (Bordeu et al. 2025) or tax policy (Brugués and DeSimone 2023). In contrast, we provide novel evidence by leveraging a natural experiment created by interest rate caps. This setting also allows us to estimate the elasticity of credit demand for corporate loans, a key parameter shaping banks' ex ante market power. Furthermore, our analysis complements recent work on market power in deposit markets (e.g., Drechsler et al. 2017, 2021 for the United States and Kulkarni and Singh 2022 for India) by documenting analogous distortions in the lending market and examining how deposit and lending market power jointly shape the impact of interest rate caps.

The rest of this paper is organized as follows. Section 2 discusses the background and describes our credit-registry data. Section 3 develops a conceptual framework and draws predictions for the effects of the interest rate cap on equilibrium credit provision. Section 4 presents our main empirical results of the effects of interest rate caps on credit provision and other outcomes. Section 5 provides additional evidence for banks' ex ante market power. Section 6 discusses policy implications. Section 7 concludes.

2 Background and Data

In this section, we provide background on the banking sector and the interest rate cap policy in Bangladesh. We then introduce our data and present descriptive evidence on the structure of the banking market before and after the cap policy was implemented.

2.1 Background

Banking Sector in Bangladesh. The banking sector in Bangladesh has undergone significant transformation since the country's independence in 1971. In the aftermath of the Bangladesh Liberation War, the government nationalized all domestic banks, aiming to direct credit toward sectors deemed vital for reconstruction. The large presence of state-owned banks under the tightly controlled market persisted until the beginning of the 1980s. This state control system was criticized for a lack of accountability and low efficiency. In the 1980s, a series of ownership reforms denationalized many of these banks. In the 1990s, further deregulatory reforms were implemented under the supervision of the World Bank, which led to entry and expansion of private-sector banks. By the 2000s, private-sector banks held more than half of the assets in this market (Asian Development Bank 2006).

As of 2008 (the beginning of our study period), there were 56 banks in Bangladesh, classified into four categories: four state-owned commercial banks (SCBs), four state-owned development financial institutions (DFIs), 39 private commercial banks (PCBs), and nine foreign commercial banks (FCBs). SCBs provide broad commercial banking services, while DFIs provide sector-specific financing, especially toward long-term development projects, agriculture, and small industries. Meanwhile, PCBs and FCBs jointly constitute the private-sector banks. As a consequence of the aforementioned deregulation reforms, PCBs and FCBs jointly constituted 62% of total assets in the banking sector as of 2008 (Moral 2016). As of 2025, there are 62 banks operating in the country—52 of which are in the private sector.

Despite the government's effort to deregulate the industry and the increased entry of private-sector banks, concerns remain about the competitiveness of the banking sector. In 2008, 43% of total assets were held by the top five banks, suggesting that the banking sector is highly concentrated (Ahamed 2012). By 2022, that percentage had fallen but was still high, at 30% (Bangladesh Bank 2022). While these numbers are similar to those reported in high-income countries, such as the United States, the banking market in Bangladesh is also characterized by a significantly smaller number of banks, raising potential concerns about banks' market power.[4]

The concentration of the banking sector raises a potential concern that banks may set lending interest rates above and beyond the costs and risks associated with disbursing loans. The high-interest-rate environment renders it difficult for firms to grow and operate at an efficient scale. As in other low-income countries, access to finance is among the most severe obstacles to firm growth reported by Bangladesh firms (Afandi and Kermani 2014).

In Bangladesh, the penetration of banking services across geography is highly heterogeneous. As we show below, bank branches are heavily concentrated in urban and economically vibrant regions, particularly in the Dhaka and Chittagong divisions, where most commercial and industrial output occurs. In contrast, rural and remote areas remain significantly underserved, with limited access to formal banking services. The uneven presence of bank branches implies that borrowers in different regions may face different interest rates, because banks may exercise varying degrees of market power across local markets due to lack of competition.

Interest Rate Caps Across the World. Interest rate caps have been widely used around the world. In a comprehensive survey, Ferrari et al. (2018) find that at least 76 countries impose some caps on lending interest rates. While these countries range across the income spectrum and stages of development, rate caps are more prevalent in low- to middle-income countries. In higher-income countries, rate cap policies are more commonly applied to individual loans (payday loans, credit cards, consumer credit) and often motivated by protecting vulnerable consumers from usurious lenders. In lower-income countries, these policies are commonly applied to a wider spectrum of loans, including corporate loans. They are often motivated as a form of industrial policy, targeting loan pricing in specific or broad sectors of the economy to expand credit and promote investment.

Despite the global prevalence of interest rate caps, they are often subject to criticisms. Some

[4]For example, as of 2024 there were 4,487 financial institutions in the United States, where the top five banks hold approximately 43% of all assets (Federal Deposit Insurance Corporation 2025).

argue these policies may discourage lending activity and induce credit rationing, especially for riskier types of borrowers. Critics also worry that these policies may encourage borrowers to take up loans beyond their repayment capacity, and thus, increase the risk of overindebtedness. Further, these policies can reduce the effectiveness of monetary policy at the macroeconomic level if lending interest rates are binding at the cap.

Interest Rate Cap in Bangladesh in 2009–2011. As a consequence of the reforms of the 1980s and 1990s, direct regulation of lending interest rates was largely phased out in Bangladesh—except for export credit, for which rates were capped at 7% in 2004. However, in the wake of the 2008 Global Financial Crisis, policy shifted amid growing concerns over slowing investment and production. On April 19, 2009, Bangladesh Bank issued a circular to mandate that commercial banks impose a maximum annualized interest rate of 13% on most business loans, citing the objective of "boosting investment" (Unnayan Onneshan 2011). The cap applied to working capital and term loans to large and medium-scale industry firms, agriculture, housing sector loans, and all types of trade financing. The capped rate was set at a high level in nominal terms, but given the high inflation rate of 7.9% and repo (repurchase) rate of 8.5% in 2008 (World Bank 2025), this cap was binding for a major percentage of loans, as we show further below.

Though the policy was introduced during a crisis, based on the lack of news coverage on the eve of the cap announcement, there is limited evidence that the market anticipated these policies. In fact, we show in our empirical results that there were limited changes in interest rates or credit provision prior to the second quarter of 2009, followed by sharp shifts immediately after the cap was enacted.

International observers, particularly the International Monetary Fund (IMF), criticized Bangladesh's rate cap policy. The IMF warned that such policies could distort credit allocation, lead to adverse consequences for financial stability, and undermine the effectiveness of monetary policy (International Monetary Fund 2011). Facing such criticisms, the central bank partially lifted the cap on corporate loans in March 2011 (for all categories except industrial term loans, preshipment credits, and agricultural sector loans) and fully in January 2012 (for industrial term loans). At the same time, still concerned about the persistence of bank market power, Bangladesh Bank introduced another regulation to cap the spread between lending and deposit rates at 100 basis points in January 2012. In this paper, we focus mainly on the introduction of the cap, rather than on its removal, because the repeal was accompanied by different policies and may confound our analysis. We leave the analysis after its partial repeal as an additional robustness analysis.

2.2 Data Sources

Our main dataset consists of confidential microdata from the Scheduled Bank Statistics (SBS), detailed reports on banks' monthly and quarterly activity (Bangladesh Bank 2013). The central bank of Bangladesh collects these data from each commercial bank for the purpose of informing monetary and banking policy. SBS consists of three components: SBS-1, which collects bank-level balance sheet and cash flow information for each month; SBS-2, which collects bank-branch-level deposit activity for each quarter; and SBS-3, which collects bank-branch-level lending activity for each quarter.

Among these three, the most important set of statistics for our purposes comes from SBS-3. For every quarter, each bank reports detailed information about outstanding loans disbursed by each branch. Banks report the number and amount of outstanding loans for each detailed bin, where a bin is defined as a combination of the branch, interest rates charged, the presence and types of collateral (e.g., real estate–secured vs. cash flow–secured), loan types, types of borrowers (i.e., corporate, government, individuals), industrial sectors for corporate borrowers, the economic purpose of the loans, and the status of overdue repayment.

While our data are not disaggregated at the individual level, they are sufficiently disaggregated that we can quantify various aspects of lending activity at the bank-branch level, such as average interest rates, the number and amount of outstanding loans, delinquency rates (overdue repayment), shares of loans secured by physical assets (hereafter referred to as "secured" loans), or the sectoral composition of loans. We can also use our data to construct measures of bank-branch-level exposure to the interest rate cap, defined by average interest rates charged above the cap in excess of the capped rate of 13% prior to the introduction of the cap. Our main empirical design uses this bank-branch-level exposure to the cap to study how the cap affected credit provision and other lending market outcomes, as we describe further in Section 4.

We complement this dataset with SBS-2, which reports the number and amount of outstanding deposits for each detailed bin, where each bin is defined as a combination of the branch, deposit rates, types of deposits (e.g., saving deposits, certificate of deposits), and types of depositors (i.e., individual or corporate). As with SBS-3, we aggregate these data at the bank-branch level, to construct the proxy for average deposit rates or deposit amount. We then merge this dataset with SBS-3 at the bank-branch and quarter level. We also merge this dataset with SBS-1, which reports bank-level balance-sheet and cash-flow information, containing detailed breakdowns of assets and liabilities; we use SBS-1 to assess heterogeneity in responses to the cap based on proxies for banks' financial health such as leverage ratio. Finally, we geocode the location of each bank branch at the subdistrict level based on banks' websites and their annual reports.

Baseline Estimation Sample. Our analysis focuses on 48 private banks: the 39 PCBs and the nine FCBs. We exclude the eight state-owned banks from our main analysis. primarily because the state-owned banks tend to have different loan objectives and offer systematically lower interest rates. By explicitly excluding these banks, we avoid our results being confounded by the differential trends of credit market outcomes between private banks and state-owned banks.

For our baseline analysis of the impacts of the interest rate cap, we use data from 2008Q1 to 2010Q4, covering five quarters before the cap was introduced (2008Q1–2009Q1) and seven quarters in the postcap period (2009Q2–2010Q4). In Appendix Figure B.5, we extend our analysis to the 2011Q1–Q4 period, after the partial removal of the cap in 2011Q1.

We focus on *corporate loans*, which we define as all loans going toward corporate entities (excluding loans toward governments, other public sectors, education, financial institutions, and individuals). We also exclude the agricultural sector from our analysis, since interest rates to firms in that sector are significantly lower than the cap throughout the sample period, meaning this sector is unlikely to be exposed by the cap.[5] Our samples include three broad sectors—manufacturing, import/export, other—covering 46 detailed sectors in total. These three broad sectors constitute approximately 46%, 17%, and 37% of precap outstanding loans, respectively (Appendix Figure A.1).

In 2008Q1, the 48 private banks had 1,917 branches with strictly positive outstanding loan balances. The number of branches increased to 2,666 by 2010Q4. In our baseline analysis, we restrict our estimation sample to a balanced panel of 1,855 branches across 39 private banks for which we observe a strictly positive amount of outstanding corporate loans in all quarters from 2008Q1 to 2010Q4.

2.3 Time-Series Evidence on Pass-Through of the Cap to Interest Rates

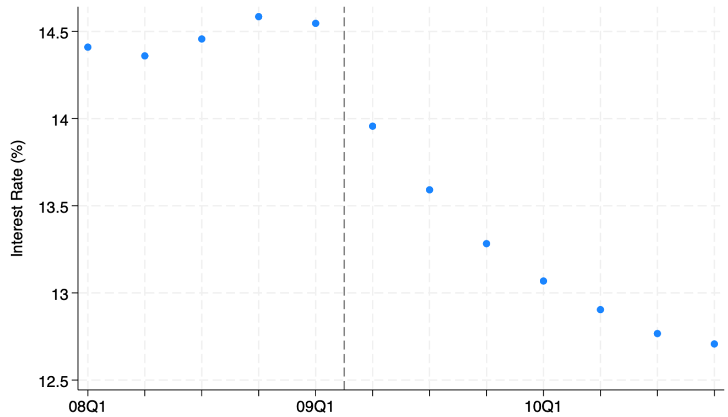
In this subsection, we report how branch-level interest rates evolved around the imposition of the cap, which motivates our branch-level DiD analysis in Section 4.

We start by providing evidence that the cap induced a sudden drop in interest rates charged on corporate loans. Figure 1 displays the transition of interest rates of corporate loans in Bangladesh over time. Panel A presents the average annualized interest rates (weighted by outstanding corporate loans) for each quarter. Prior to the introduction of the cap in 2009Q2, interest rates hovered around 14.5%, without any clear time trend. As soon as the cap was introduced, average interest rates suddenly decreased. Average rates continued to fall, reaching below 13% in 2010Q2.

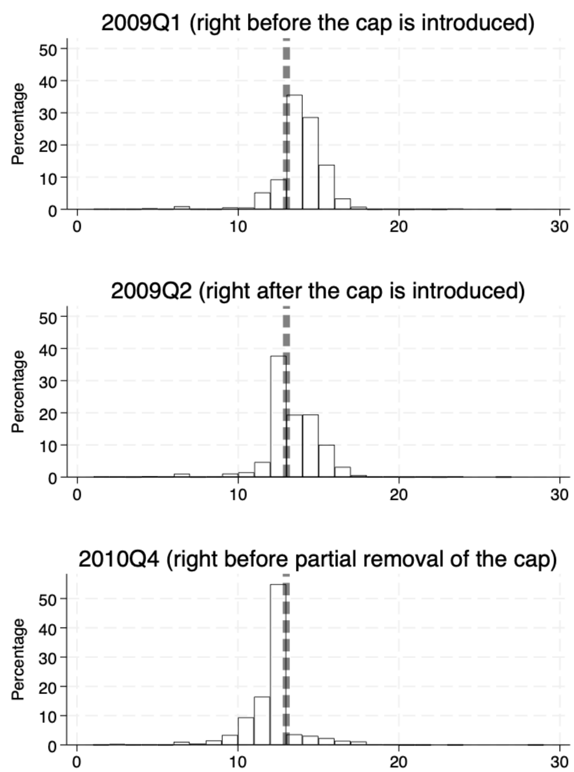
[5]While the cap applied only to "large and medium" firms, we include all types of loans (regardless of firm size) because we do not observe a precise firm-size classification. While this may lead to partial compliance for the reduced-form effects of caps on interest rate changes, our instrumental variable (IV) effects of the cap-induced changes in rates on credit outcomes remain valid when interpreted as local average treatment effects (LATE).

Figure 1: Evolution of Interest Rates for Corporate Loans Around the Cap Policy

Panel A: Average Interest Rates



Panel B: Distribution of Outstanding Loans by Interest Rate (%)



Notes: Panel A presents the evolution of average interest rates on corporate loans, weighted by outstanding loan amount shares, using our baseline SBS-3 data. The dotted line marks the timing of the interest rate cap's introduction, shortly before the start of 2009Q2. Panel B displays the distribution of outstanding loans across interest rate bins for three periods: right before the cap was introduced (2009Q1), right after its implementation (2009Q2), and shortly before its partial repeal by Bangladesh Bank (2010Q4). The dotted line indicates 13%, the rate ceiling imposed by the cap.

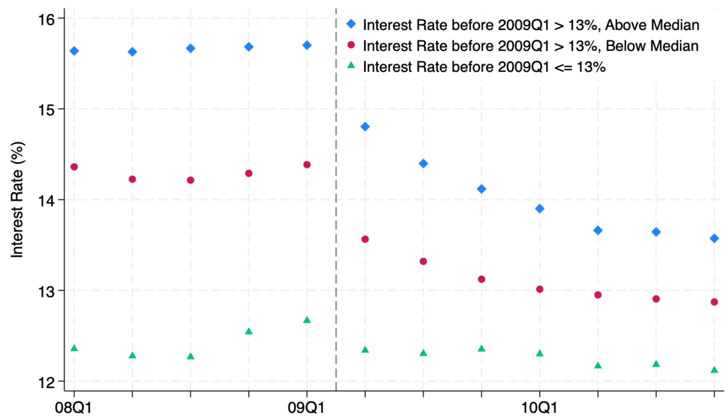
They did not fall to below 13% right after the cap's implementation because the policy did not apply retroactively to loans originated before the cap was introduced. As time went on, previously disbursed loans reached their maturity, and their share in overall outstanding loans decreased. The limited time trend of the interest rate up to 2009Q1, and its sudden drop in 2009Q2, jointly support the interpretation that the interest rate cap was not anticipated by the market.

In Panel B, we present further evidence of the market-wide effects of the cap by displaying the evolution of the distribution of outstanding loans by interest rate charged. Panel B displays the distribution of outstanding loans across interest rate bins for three periods: right before the cap was introduced (2009Q1), right after its implementation (2009Q2), and shortly before its partial repeal (2010Q4). Before the cap was introduced, a majority of outstanding loans were charging interest rates above 13%. Right after the cap was introduced, the mass of loans charging interest rates of exactly 13% suddenly increased. Again, because the interest cap did not apply retroactively, there was still a nontrivial percentage of outstanding loans with interest rates above 13%. By 2010Q4, this share had shrunk, while the mass of loans at the 13% threshold continued to increase. The presence of bunching at 13% interest rates suggests that the cap was binding for a large fraction of outstanding loans and had a tangible impact on the corporate loan market.[6]

We now turn to how the cap influenced the interest rates charged across different bank branches. In Figure 2, we plot the trajectory of average interest rates across branches, which we sort into three groups based on their precap rates: (i) those below 13%, (ii) those above 13% but below the median within this sample, and (iii) those above 13% and above the median within this sample. Before the cap was introduced, the interest rates of all three types of branches remained relatively flat, again suggesting that there was limited anticipation of the policy by the market. Once the cap was introduced, branches with higher precap interest rates sharply reduced their interest rates relative to those with lower precap rates. In contrast, branches whose interest rates

[6] In Appendix Figure A.2, we plot the transition of the share of loans with annualized interest rates equal to or below 13% over time; we show that this share suddenly increased after 2009Q2, consistent with these figures.

Figure 2: Interest Rates Across Bank Branches by Cap Exposure



Notes: This figure plots the path of average interest rates (weighted by outstanding loan-amount shares) dividing branches into three strata of cap exposure based on precap average interest rates: (i) those below 13%, (ii) those above 13% but below the median within this sample, and (iii) those above 13% and above the median within this sample.

were below 13% barely changed their interest rates. Consequently, the gap in interest rates across these three strata suddenly narrowed after the cap.

Figure 2 indicates that the effects of interest rate caps may have differed substantially across bank branches depending on their prevailing preregulation interest rates. In Section 4, we use this variation in preregulation branch-level interest rates to inform a DiD research design, which we apply to study the effects of interest rate caps on equilibrium credit provision.

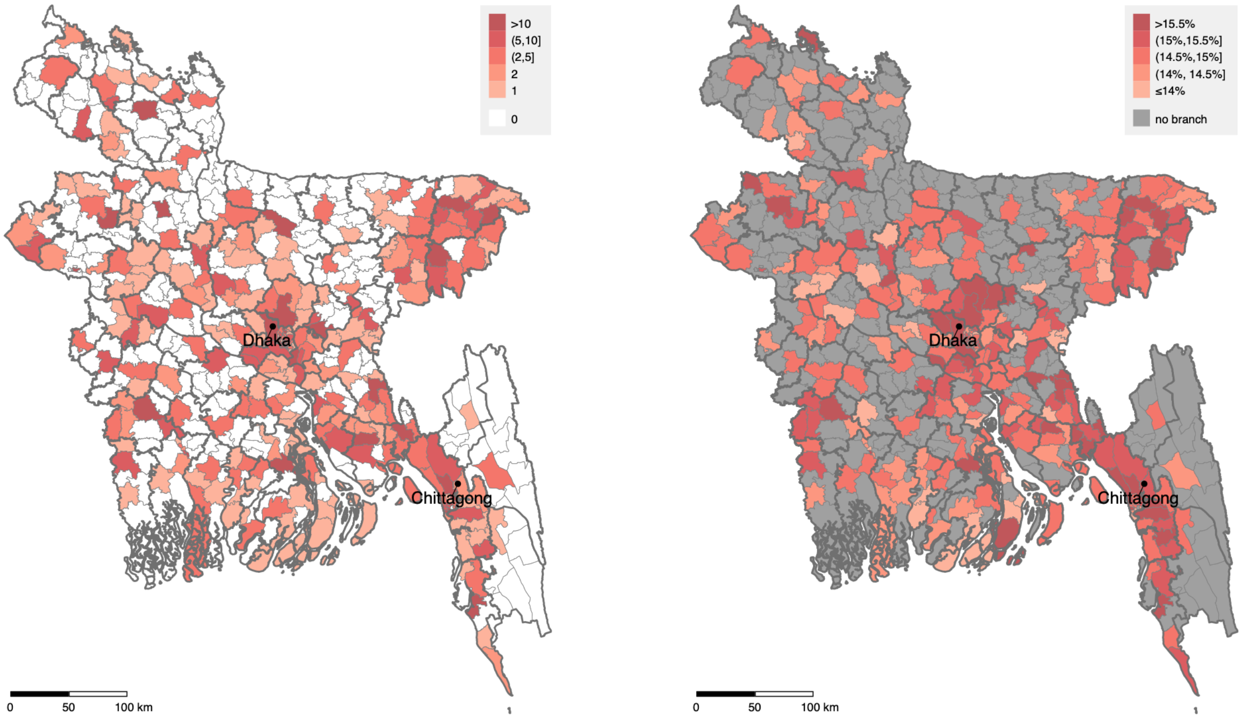
2.4 Precap Interest Rates and Credit Provision Across Bank Branches

In this subsection, we document the heterogeneity of interest rates and key credit market characteristics across branches prior to the cap (2008Q1–2009Q1).

In Figure 3, we show the spatial distribution of bank branches and average interest rates charged. Bangladesh is administratively divided into two layers: 64 districts (*zila*) and 495 subdistricts (*upazila*). In this figure, darker gray lines indicate district borders, while lighter ones indicate subdistrict borders. Each subdistrict consists of an average geographic area of 300 square

Figure 3: Bank Branches and Interest Rates Across Subdistricts in 2008 (Precap)

Panel A: Number of Branches **Panel B: Average Interest Rates**



Notes: This figure displays the precap distribution across subdistricts (*upazila*) of bank branches in Panel A and average share-weighted interest rates charged on business loans in Panel B. There are 495 subdistricts (thinner lines) nested within 64 districts (thicker lines) in Bangladesh. In Panel A, we display subdistricts with no branches (in white). In Panel B, we display subdistricts with no business loans (in gray), and therefore no interest rate data.

kilometers, which is roughly one-sixth the size of the average U.S. county.

Panel A shows the number of bank branches for each subdistrict among the 48 private banks in our sample. We observe significant heterogeneity in the number of bank branches. Bank branches are more concentrated around the Dhaka and Chittagong districts, which are the economic centers of Bangladesh. In Panel B, we present the average interest rate charged for each subdistrict (weighted by outstanding loan-amount shares), aggregating all bank branches present in those districts. While average precap interest rates across all branches were around 14.5% (Figure 1), there is significant spatial heterogeneity across these subdistricts.[7]

[7]See Appendix Figure A.4 for the histogram of precap rates across branches.

Table 1: Patterns of Precap Branch-Level Characteristics

	Interest Rate	log Number of Outstanding Loans	log Average Outstanding Amount	Deposit Rate	Delinquency Rate: $\geq$ 9 Months	Proportion of Secured Loans
	(1)	(2)	(3)	(4)	(5)	(6)
ln(Population Density)	-0.100***	0.101***	0.376***	0.000	0.005***	-0.009**
	(0.009)	(0.015)	(0.014)	(0.001)	(0.001)	(0.004)
Subdistrict-Bank HHI	0.242***	-0.299***	-0.612***	0.009	-0.014	0.007
	(0.053)	(0.109)	(0.095)	(0.010)	(0.010)	(0.027)
Specification	OLS	OLS	OLS	OLS	OLS	OLS
Bank FE	X	X	X	X	X	X
Quarter FE	X	X	X	X	X	X
Mean	14.938	4.161	2.575	5.901	0.044	0.623
Number of Banks	39	39	39	39	39	39
Number of Branches	2124	2124	2124	2124	2124	2124
Observations	9959	9959	9959	9959	9959	9959
Adj. R-squared	0.559	0.392	0.536	0.885	0.214	0.273

Notes: Each column in this table reports results from estimating a predictive regression of precap branch-level average corporate-loan interest rates on branch characteristics, controlling for bank and quarter fixed effects. We compute HHIs based on outstanding loan amounts at the subdistrict level. *Average Outstanding Amount* in column (3) is defined as the total outstanding loan amount divided by the number of loan contracts with a remaining balance. *Secured Loans* in column (6) refers to loans for which the borrower has pledged real estate or machinery as collateral.

In Table 1, we further investigate the patterns of heterogeneity in interest rates and other credit activity across branches within the same bank and quarter. Specifically, we run a regression of branch-level credit market characteristics on the log of population density and the HHI in terms of outstanding loans of the bank within the subdistrict.[8] We control for bank fixed effects to isolate how credit provision systematically differs across different branches based on the local market characteristics.

In column (1), we find that average interest rates are lower in subdistricts with higher population density and lower HHI. These results are consistent with the interpretation that banks may have market power to set interest rates. However, they are also compatible with alternative explanations, such as lower costs of credit provision in more densely populated or less competitive areas. In Section 4, we leverage the natural experiment of the introduction of the interest rate cap to isolate these alternative channels.[9]

[8] If a bank has multiple branches in the same subdistrict, we consider them as one entity when we compute the HHI.

[9] In Appendix Table A.1, we show that these patterns hold true even after simultaneously controlling for deposit rates, delinquency rates, and the proportion of secured loans (column 4).

In columns (2) and (3), we find that the number of outstanding loans and the average outstanding amount are higher in subdistricts with higher population density and lower HHI. The relation with population density is consistent with the interpretation that there is a greater demand for credit in more populated areas. The patterns for HHI can be explained by the fact that banks tend to open branches in areas with high credit demand, leading to a negative association between HHI and credit provision.

In column (4), we find that deposit rates are not statistically associated with population density or the HHI. Notably, the magnitudes of these coefficients are also close to zero, especially when compared to those for lending rates in column (1). This suggests that, in this context, variation in lending rates across branches is a more important driver of interest rate spreads than variation in deposit rates.[10]

Columns (5) and (6) investigate the patterns of borrower risk across branches. In column (5), we report regression results for delinquency rates, defined by the percentage of loans where the repayment is overdue by more than nine months.[11] We find that delinquency rates are positively associated with population density. In column (6), we find that the proportion of secured loans—defined as loans that are associated with physical assets (i.e., machinery or real estate) used as collateral—declines with population density. Lenders often require physical assets to secure loans to incentivize repayment for risky borrowers (Berger and Udell 1995). Therefore, higher delinquency rates and a lower share of secured loans in more populous areas may reflect greater difficulty banks face in having borrowers pledge collateral in those areas, resulting in higher equilibrium delinquency rates.

[10]Appendix Table A.2 replicates the analysis using the HHI based on deposit amounts rather than lending amounts. The estimated coefficient is slightly larger (0.018) with a standard error of 0.011, aligning with prior findings on deposit market power in the United States (e.g., Drechsler et al. 2017, 2021). Nevertheless, the variation remains smaller than that observed for lending rates. In Section 4.4, we further examine the branch-level impact of lending rate caps on deposit rates and find no statistically significant effects.

[11]Loans with payments overdue by more than nine months are typically classified as non-performing loans (NPLs) in World Bank member countries such as Bangladesh (Adhikary 2006). The patterns are nearly identical if we instead use three-month or six-month delinquency thresholds.

3 Conceptual Framework

In this section, we develop a simple model of credit markets with imperfectly competitive banks to study the effects of interest rate caps on corporate lending. We use this model to highlight which economic forces and parameters shape the effects of interest rate caps on equilibrium credit provision.

3.1 Model

There is a continuum of entrepreneurs of measure 1 seeking financing for projects. At date 1, each borrower has a project that requires one unit of goods as an investment. Hereafter, we take these goods as a numéraire and normalize the price to one.[12] If invested, the project succeeds and yields return R_1 with probability p, and fails by offering a zero return with probability $1-p$ at the end of date 1. Furthermore, if the project is successful in date 1, the entrepreneur will start another project at the beginning of date 2. This project requires one unit of goods as an investment at the beginning of date 2 and returns R_2 with probability 1 at the end of date 2. If the first project is not financed on date 1, the borrower does not have an investment opportunity in date 2.

There is no storage technology in this economy, and hence, borrowers need to finance their projects through banks on both dates. For expositional purposes, we assume that borrowers fully discount the future and maximize profits period by period.[13] We assume that there is a representative bank that operates under imperfect competition. The bank's cost of raising one unit of funding is c in both periods.

On date 1, the bank posts the interest rate on new loans r_1. Based on this interest rate,

[12]Here, we abstract from inflation, and hence all interest rates are in real terms. In Section 4, we convert nominal interest rates to real rates when we compute credit demand elasticities implied by our estimates.

[13]We assume a forward-looking borrower does not alter any equilibrium outcome in the absence of interest rate caps, as banks can in any case extract the entire surplus in the second period. If the cap binds in the second period, the expression in Proposition 1 is slightly modified in this extension, since borrowers anticipate retaining some surplus if they borrow in date 1. Nonetheless, our main conclusion regarding the roles of ex ante and ex post market power remains unchanged.

each entrepreneur determines whether to borrow from the bank and invest in the project. To borrow from the bank, each entrepreneur i incurs a fixed cost u_i, capturing various idiosyncratic factors associated with entrepreneurs' access to banks. u_i is distributed following the cumulative distribution function $F(\cdot)$, which has full support and is twice continuously differentiable. If the project succeeds, the entrepreneur earns a return of R_1 and repays r_1 to the bank. Otherwise, the project yields no return, and the entrepreneur does not repay the loan.

At the beginning of date 2, the bank observes which entrepreneurs succeeded with their date 1 projects and thus have an investment opportunity in date 2. We assume that the bank can lock in such borrowers with probability γ. In this event, the bank can fully extract rents from these borrowers by charging an interest rate R_2. With the remaining probability $(1-\gamma)$, borrowers leave the bank and seek funding through other means at the interest rate R_2. Therefore, the parameter γ proxies banks' ex post market power after establishing a relationship with the borrower.

On date 1, the bank sets interest rates r_1 to maximize its expected profit. An entrepreneur decides to borrow on date 1 if the expected profit exceeds the fixed cost, where the expected profit is the product of their project surplus minus the repayment, multiplied by the project's success probability. Specifically, entrepreneur i with cost u_i borrows from the bank if and only if

$$p(R_1 - r_1) - u_i > 0 \tag{1}$$

By integrating over u_i, the demand function is given by $D(r_1) = F(p(R_1 - r_1))$. We denote the elasticity of demand by $\epsilon(r_1) \equiv -\frac{r_1}{D(r_1)} \frac{\partial D(r_1)}{\partial r_1}$.

To define the equilibrium interest rate, we first consider the break-even interest rate for the bank in date 1. For each entrepreneur, the cost of funding on both dates is c. The bank breaks even if this cost equals the expected benefit in date 1, pr_1, plus the expected profit in date 2, $p\gamma(R_2 - c)$. Hence, the break-even interest rate for the bank is $c/p - \gamma(R_2 - c)$. Following the approach of Weyl and Fabinger (2013), we introduce imperfect competition to this setting using a conduct parameter. Namely, we assume that the bank charges interest rates as follows:

$$r_1 = \frac{c/p - \gamma(R_2 - c)}{1 - \theta/\epsilon(r_1)} \tag{2}$$

where θ is the conduct parameter that governs the degree of imperfect competition. If $\theta = 1$, the bank operates as a monopoly, and when $\theta = 0$, the market is under perfect competition.

3.2 Predicted Effects of an Interest Rate Cap

We now analyze the impacts of an interest rate cap policy in this lending market with imperfect competition. Consider the interest rate cap in dates 1 and 2 at $\bar{r} = \delta r_1$, where $0 < \delta < 1$. To focus on a nontrivial case, assume that $R_2 > \bar{r}$, so that the interest rate cap also binds in date 2. Because the bank's surplus goes down in date 2, the bank's break-even interest rate on date 1 goes up to

$$c/p - \gamma(\bar{r} - c) \tag{3}$$

If this break-even rate is still below the interest rate cap, $\bar{r}$, banks keep supplying credit at interest rate $\bar{r}$. In this case, lower interest rates attract credit demand and increase equilibrium credit provision. If the break-even rate is above the cap, the bank is unwilling to supply credit because its expected profit is zero. We summarize this result via the following proposition:

Proposition 1 *Consider the policy to cap the interest rate at $\bar{r} = \delta r_1$ in dates 1 and 2, where $0 < \delta < 1$ and $R_2 > \bar{r}$. Equilibrium credit provision in date 1 strictly increases if*

$$\delta > \left(1 - \frac{\theta}{\epsilon(r_1)}\right) \frac{c/p - \gamma(\bar{r} - c)}{c/p - \gamma(R_2 - c)} \tag{4}$$

and the increase in credit provision in date 1 is given by $D(\bar{r}_1) - D(r_1)$, i.e., it is purely determined by demand. Otherwise, equilibrium credit provision is zero (credit rationing).

Proposition 1 clarifies under what conditions the interest rate cap leads to increases or decreases in equilibrium credit provision. In particular, it clarifies how ex ante market power (proxied by θ) and ex post market power (proxied by γ) shape the cap's impacts. First, this condition is more likely to be satisfied if ex ante market power, θ, is greater. In particular, if $\theta = 0$ (i.e., perfect competition), the equilibrium interest rate coincides with the break-even interest rate (equation 2) even without regulation. In this case, any binding $\bar{r}$ results in a decrease in credit provision. Second, this condition is less likely to be satisfied if γ is larger. Intuitively, if banks supply credit in date 1 under the assumption that they can extract more surplus in date 2 (a higher γ), the binding interest rate in date 2 limits the scope for surplus extraction and decreases expected profit.

Another important implication of this proposition is that, as long as the condition in Proposition 1 holds, the increase in equilibrium credit provision is purely determined by demand, and it does not directly depend on factors such as funding costs c or repayment rates. Indeed, in this case, the observed effects of interest rate cap allow us to estimate the credit demand elasticity $(\log D(\bar{r}_1) - \log D(r_1)) \,/\, (\log \bar{r}_1 - \log r_1) \approx -\epsilon(r_1)$. We test these predictions and estimate credit demand elasticities in Section 4.

Finally, it is straightforward to see that setting the interest rate cap at the bank's break-even interest rate maximizes consumers' and lenders' joint surplus. Therefore, if the condition in Proposition 1 is satisfied, the interest cap is welfare-improving. This welfare improvement arises because of the expansion in credit provision in both date 1 and date 2. As mentioned above, this condition is more likely to be satisfied if ex ante market power (θ) is stronger and ex post market power (γ) is weaker.

The model presented here is a partial equilibrium model, and hence omits various general equilibrium effects, such as the effects of the cap on the cost of funding c. We also abstract from other policy considerations, such as the effectiveness of monetary policy in the presence of the cap. The purpose of the model is not to draw broader conclusions about the effectiveness of the cap policy on aggregate welfare; rather, the purpose is to illustrate how banks' ex ante and ex post market power jointly shape the effects of interest rate caps on credit provision, which we empirically examine in the following sections.

4 Main Analysis: Credit Market Effects of the Interest Rate Cap

In this section, we present our empirical results on the effects of the interest rate cap in Bangladesh from 2009 to 2011 on branch-level credit provision and other market outcomes.

4.1 Empirical Specifications

Identifying the causal effects of the interest rate cap presents an empirical challenge, as the policy was implemented uniformly for most categories of corporate loans all throughout Bangladesh, as described in Section 2.1. Although our dataset spans the pre- and postcap periods, simply tracing out aggregate credit provision over time is not informative about the impact of the cap, as it is confounded by broader macroeconomic trends.

To address this challenge, we implement a DiD strategy that leverages branch-level variation in precap interest rates. As discussed in Section 2.3, the introduction of the cap led to sharp declines in lending rates for branches that were charging rates above 13% prior to the policy. In contrast, branches with precap rates below 13% experienced little to no change in their lending rates. This heterogeneity in exposure allows us to identify the cap's effects on credit market outcomes by comparing branches more vs. less affected by the policy.

Specifically, we estimate the following event-study regression specification:

$$Y_{i,t} = \sum_{s=-m,s\neq-1}^{s=n} \beta_s \cdot TrtIntensity_i \times \mathbb{1}\{t=s\} + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \epsilon_{i,t} \quad (5)$$

where i is the bank branch, t is the quarter, $Bank(i)$ is the bank that branch i belongs to,

$District(i)$ is the district that branch i belongs to, and $Y_{i,t}$ is the outcome variable (e.g., interest rates, amount and number of outstanding loans, fraction of delinquent loans). We normalize the quarter such that $t = 0$ corresponds to 2009Q2, the quarter right after the introduction of the cap. We set the beginning of the sample at $m = -5$ (2008Q1) and the end of the sample at $n = +7$ (2010Q4).

$TrtIntensity_i$, which captures how much branch i is exposed to the interest rate cap regulation, is constructed as follows: we first compute the average of annualized interest rates (weighted by outstanding loans) on outstanding loans by bank branch i during the precap period (2008Q1–2019Q1). If that average is above 13%, we then take the difference between the branch-level average interest rate and the 13% cap. If it is below 13%, we assign $TrtIntensity_i = 0$ to the branch.[14] We include branch fixed effects η_i, which ensures that β_s are identified off of the comparison between the changes in outcome variables over time. We omit the quarter before the reform, β_s for $s = 2009Q1$, which serves as the reference period for the event-study coefficients. Hence, β_s captures the marginal increase in the changes of the outcome variable in quarter t relative to the reference period (2009Q1) if the preregulation interest rate increases by one percentage point.

We also include bank-quarter fixed effects $\nu_{Bank(i),t}$ in all specifications. These fixed effects allow us to account for time-varying shocks to banks, including any bank-specific fallout from the Global Financial Crisis or bank-specific changes to capital requirements accompanying the adoption of Basel II regulations in Bangladesh in 2010 (Bangladesh Bank 2008). Furthermore, we also include district-quarter fixed effects $\psi_{District(i),t}$ to rule out differential changes in regional economic activity, such as faster growth in industrial activity in the Dhaka or Chittagong districts.

Our identifying assumption is parallel trends of potential outcomes across branches with different treatment intensity ($TrtIntensity_i$) in the absence of the cap, after controlling for the

[14]As a robustness check, we present results from estimating Equation (5) using a discretized version of $TrtIntensity_i$ in Figure B.3, and we find that the estimated effects on branch-level lending outcomes are monotonic in quantiles of treatment intensity.

time-varying bank-level and district-level changes in outcome variables. As a check on this identifying assumption, we test whether β_s is insignificant and close to zero before the regulation starts. Furthermore, as a robustness exercise, we present our results by controlling for the precap growth of the interest rate (from 2008Q1 to 2009Q1) interacted with quarter dummies. As we show below, controlling for precap rate trends barely changes our results, consistent with our finding in Figure 2 that the preperiod growth in interest rates is uncorrelated with our proxy for branch-level treatment intensity ($TrtIntensity_i$).

Although regression equation (5) is informative about differential impacts of the interest rate cap regulation across bank branches, it is also useful to convert the reduced-form estimates to determine how much a 100-basis-point change in interest rates from the cap leads to a change in credit market outcomes. To answer this question, we estimate the following IV regression:

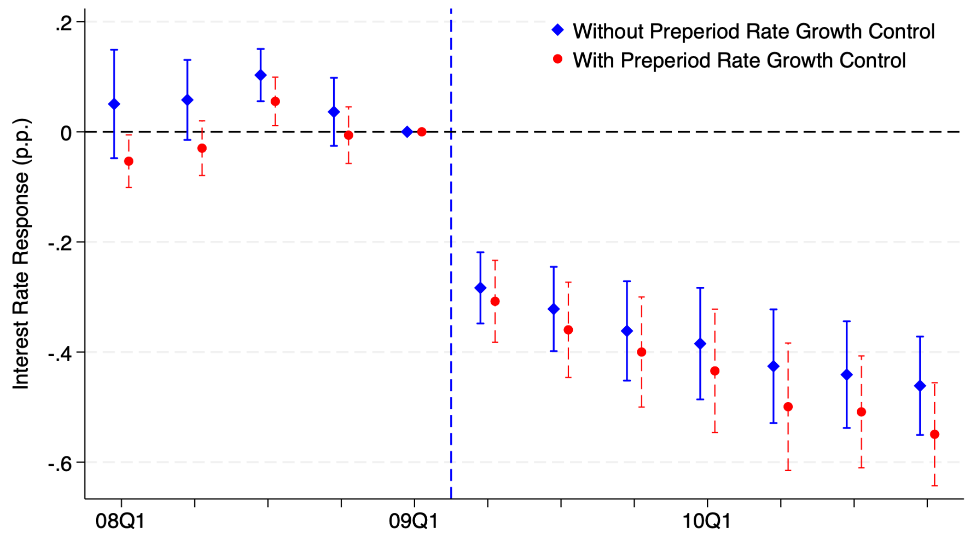
$$Y_{i,t} = \alpha \cdot InterestRate_{i,t} + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \epsilon_{i,t} \tag{6}$$

where we instrument $InterestRate_{i,t}$ with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$. As in specification (5), we also include branch fixed effects η_i, bank-quarter fixed effects $\nu_{Bank(i),t}$, and district-quarter fixed effects $\psi_{District(i),t}$. This instrument extrapolates the effects of the decrease in interest rates on outcome variables using the change in interest rates induced by the introduction of the cap. When the outcome variable $Y_{i,t}$ proxies log credit provision, the resulting IV estimate of α thus corresponds to an interest rate semielasticity. From the perspective of our model in Section 3, if banks have ex ante market power and the cap is still above their break-even rates, this semielasticity corresponds to the semielasticity of credit *demand*.

4.2 Effects on Interest Rates

We first verify that bank branches with a greater treatment intensity indeed face a sudden, faster decline in interest rates. In Figure 4, we present the estimates of $\widehat{\beta}_s$ from regression (5)

Figure 4: Event-Study Effects of the Rate Cap on Interest Rates



Notes: This figure plots the estimated coefficients of treatment intensity interacted with quarter dummies from event-study regression (5) with the share-weighted average branch-level interest rate charged on corporate loans as the outcome. We omit the quarter before the cap reform (2009Q1) as the reference category. The two series of reported coefficients correspond to the specification with or without controlling for the precap growth of the interest rate (from 2008Q1 to 2009Q1) interacted with quarter dummies. The vertical line indicates the timing of the introduction of the interest rate cap (April 19, 2009; 2009Q2). 95% confidence interval bars are obtained from clustering standard errors at the branch level.

with branch-level average interest rates as the outcome. The two series of reported coefficients correspond to our baseline specification ("without preperiod rate-growth control") and the specification by controlling for the precap growth of the interest rate (from 2008Q1 to 2009Q1) interacted with quarter dummies ("with preperiod rate-growth control").

As anticipated from Figure 2, there are strong, sudden negative effects on branch-level interest rates. Before the regulation, there is no discernible pretrend for our baseline specification. Right after the central bank introduced the cap, interest rates immediately declined, indicating that bank branches with high treatment intensity (i.e., those that ex ante charged rates above 13%) decreased their interest rates by relatively more. Average interest rates responded gradually to the introduction of the cap, primarily because some outstanding loans were disbursed before the cap was introduced, and thus were not subject to the cap policy. In response to a one-percentage-point increase in treatment intensity—that is, a 100-basis-point spread between prereform interest rates and the 13% cap—average interest rates decline by approximately 30 basis points in 2009Q2, and the effect grows to an approximately 50-basis-point decline by 2010Q4. We find virtually identical results for the specification with preperiod rate-growth control, consistent with the observation that the preperiod growth in interest rates is not correlated with our proxy for the treatment intensity ($TrtIntensity_i$).

4.3 Average Effects of the Cap on Credit Provision

The preceding set of results indicates that bank branches that were more exposed to the cap sharply lowered interest rates after the cap was introduced. In this subsection, we study how this branch-level exposure to the lending rate cap affected credit provision.

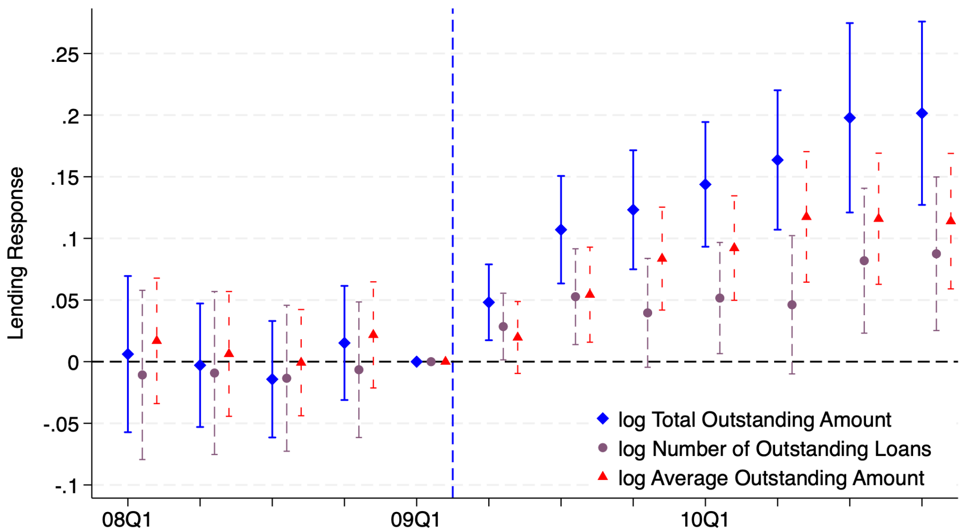
Baseline Event-Study Results. Figure 5 shows the effects on branch-level log total outstanding loan amounts, the total number of outstanding loans, and the average outstanding amount per loan account (excluding loans whose repayment was more than nine months past due). There are no statistically significant pretrends prior to the regulation, bolstering the validity of our DiD design.

Right after the cap was introduced, it led to an *increase* in corporate credit provision for branches with a high treatment intensity, with the effect growing over the course of the cap regime. The effect on total outstanding loan amounts gradually increased from 5.1% (0.05 log points) in 2009Q2 to 22.1% (0.20 log points) by 2010Q4. This increase is driven by both the extensive margin (an increase in the number of loans) and the intensive margin (an increase in the average loan size); the latter made a slightly larger contribution.[15]

Table 2 summarizes these results in a regression table format by pooling coefficients across quarters. Columns (1) and (2) show the coefficients for the effect on interest rates; the remaining

[15] The patterns are nearly identical if we control for preperiod interest rate growth, interacted with quarter fixed effects (Appendix Figure B.1).

Figure 5: Event-Study Effects of the Rate Cap on Credit Provision



Notes: This figure plots the estimated coefficients of treatment intensity interacted with quarter dummies from event-study regression (5) with corporate lending measures as the outcome. We consider three measures of equilibrium branch-level credit provision: the log of total outstanding loan dollars, the log number of outstanding loans (extensive margin), and the log average outstanding amount, computed as total lending dollars divided by the number of loans (intensive margin). We omit the quarter before the cap reform (2009Q1) as the reference category. The vertical line indicates the introduction of the interest rate cap (April 19, 2009; 2009Q2). 95% confidence interval bars are obtained from clustering standard errors at the branch level.

columns use a measure of credit provision as the outcome. Mirroring Figure 5, our results indicate that banks responded to the cap by expanding their lending on both the intensive and extensive margins. Our results hold regardless of whether we exclude the preperiod rate-growth controls (odd columns) or include them (even columns).[16]

Through the lens of the model in Section 3, our finding is consistent with the interpretation that banks wield ex ante market power, and that the cap did not induce credit rationing. Our empirical findings contrast sharply with other studies on lending rate caps focusing on high- or upper-middle-income countries and those imposed on loans to consumers or small firms. By implementing a similar DiD design, existing work in those settings typically finds a *decline* in

[16]The point estimates are nearly identical if we account for finer geographic trends by including subdistrict-quarter fixed effects (Appendix Figure B.2).

Table 2: Pooled Event-Study Effects of the Rate Cap on Interest Rates and Credit Provision

	Interest Rate		log Total Outstanding Amount		log Number of Outstanding Loans		log Average Outstanding Amount	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
TrtIntensity x 2008Q1-Q4	0.06*	-0.01	0.00	0.01	-0.01	-0.01	0.01	0.01
	(0.03)	(0.01)	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)
TrtIntensity x 2009Q2-Q4	-0.33***	-0.36***	0.09***	0.10***	0.04**	0.05***	0.05***	0.05***
	(0.04)	(0.04)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
TrtIntensity x 2010Q1-Q4	-0.46***	-0.50***	0.18***	0.18***	0.07**	0.07***	0.11***	0.11***
	(0.05)	(0.05)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)
Specification	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
Branch FE	X	X	X	X	X	X	X	X
Bank X Quarter FE	X	X	X	X	X	X	X	X
District X Quarter FE	X	X	X	X	X	X	X	X
Preperiod Rate Growth Control		X		X		X		X
Baseline Mean	14.90	14.90	6.91	6.91	4.33	4.33	2.58	2.58
Number of Banks	39	39	39	39	39	39	39	39
Number of Branches	1855	1855	1855	1855	1855	1855	1855	1855
Observations	22260	22260	22260	22260	22260	22260	22260	22260
Adj. R-squared	0.87	0.88	0.95	0.95	0.91	0.91	0.92	0.92

Notes: This table reports results from estimating event-study regression (5) with effects pooled across several quarters. $TrtIntensity_i$ captures how much branch *i* is exposed to the interest rate cap regulation; it is constructed as follows: we first take average annualized interest rates of outstanding loans by bank branch *i* from the first quarter of 2008 to the first quarter of 2009. If this number is above 13%, we take the difference between the average interest rate and the 13% cap threshold. If the difference is below 13%, we assign $TrtIntensity_i = 0$, indicating that the branch is, on average, inframarginal to the reform. All estimates are relative to the quarter before the reform, and we therefore omit $TrtIntensity \times 2009Q1$. We define the interest rate outcome in columns (1) and (2) as the share-weighted average branch-level interest rate on corporate loans. Odd columns correspond to our baseline specification; in even columns, we include preperiod rate-growth controls. Robust standard errors clustered at the branch level are in parentheses. *p<0.1; **p<0.05; ***p<0.01.

credit provision for consumer or small-firm loans, as we discussed in Section 1. This divergence may reflect the distinct institutional setting we study: a low-income country (Bangladesh), where capital markets are less developed, and our focus on large corporate loans, where only a limited number banks can offer loans due to scale and are therefore potentially more likely to exercise substantial market power.

IV Effects of Cap-Induced Rate Changes. We next estimate the IV regression (6). This specification translates the reduced-form estimates reported in Table 2 into a semielasticity of credit provision with respect to the interest rate. If the cap lies above the break-even interest rate, as suggested by the positive reduced-form effects of the cap on credit provision above, this semielasticity can be interpreted as the semielasticity of credit demand, as implied by our model in Section 3.

Table 3 presents our results. Across the board, we compute large first-stage Montiel Olea and Pflueger (2013) F-statistics of above 100, and exceed the two-stage least squares (TSLS) estimator thresholds for 5% worst-case bias relative to OLS at the 5% confidence level, consistent with sharp changes in interest rates induced by the cap (Figure 4). In column (1), we find that a one-percentage-point decrease in average interest rates increases branch-level total outstanding loan dollars by 31 log points ($\approx$ 36%). These effects are driven by both the 14-log-point ($\approx$ 15%) increase in extensive-margin lending (number of outstanding loans) in column (3), and the 17-log-point ($\approx$ 19%) increase in intensive-margin (average outstanding amount per loan account) lending in column (5). Our results hold when we adopt our baseline specification (odd columns), or the specification where we include the preperiod rate-growth controls (even columns).

For ease of interpretation, we convert the estimated semielasticity into an elasticity of credit demand with respect to the real interest rate. Consider a loan whose nominal annualized interest rate decreases from 14% (slightly less than the precap average of 14.5%; Figure 2) to the rate ceiling of 13%. The corresponding elasticity is approximately $1.7 \approx -0.31/\log\left(\frac{0.14-0.079}{0.13-0.079}\right)$, using the 2008 inflation rate of 7.9% reported by the World Bank (Section 2.1). As mentioned above, slightly less than half of the total lending response arises from the extensive margin, with the remainder driven by the intensive margin.

We are not aware of existing studies that provide a credible estimate of the elasticity of credit demand for large corporate loans in a developing country. This is likely due to limited data availability and challenges in finding exogenous variation in interest rates. For smaller-scale loans in the microfinance sector, Karlan and Zinman (2008) estimate an extensive-margin demand elasticity of 0.28 in a randomized control trial in South Africa. Dehejia et al. (2012) exploit quasi-experimental variation in loan pricing by a microfinance institution in Bangladesh and

Table 3: IV Estimates of Cap-Induced Changes in Interest Rates on Credit Provision

	log Total Outstanding Amount		log Number of Outstanding Loans		log Average Outstanding Amount	
	(1)	(2)	(3)	(4)	(5)	(6)
Interest Rate	-0.31***	-0.33***	-0.14**	-0.16***	-0.17***	-0.18***
	(0.06)	(0.06)	(0.06)	(0.06)	(0.05)	(0.05)
Specification	IV	IV	IV	IV	IV	IV
Montiel Olea and Pflueger F-Statistics	109.54	106.07	109.54	106.07	109.54	106.07
TSLS 5% Critical Value	37.42	37.42	37.42	37.42	37.42	37.42
Branch FE	X	X	X	X	X	X
Bank X Quarter FE	X	X	X	X	X	X
District X Quarter FE	X	X	X	X	X	X
Preperiod Rate Growth Control		X		X		X
Baseline Mean	6.91	6.91	4.33	4.33	2.58	2.58
Number of Banks	39	39	39	39	39	39
Number of Branches	1855	1855	1855	1855	1855	1855
Observations	22260	22260	22260	22260	22260	22260

Notes: This table reports results from estimating the IV Regresion (6) via TSLS. We instrument the endogenous variable, the branch-level average interest rate, with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Odd columns correspond to our baseline specification, in which we include bank-quarter fixed effects and district-quarter fixed effects; in even columns, we control for the preperiod growth rate in interest rates. Robust standard errors clustered at the branch level are in parentheses. For each specification, we report the first-stage F-statistic for the excluded instrument from Montiel Olea and Pflueger (2013), which is robust to clustering by bank branch and to heteroskedasticity. We also report the 5% critical-value thresholds for 5% worst-case bias in the TSLS estimates relative to OLS. *p<0.1; **p<0.05; ***p<0.01.

estimate overall elasticities ranging from 0.73 to 1.04. Our estimated elasticity is higher than these studies. This is consistent with the view that larger firms face more elastic demand due to having access to alternative sources of financing, such as bond markets. In the context of advanced economies, Altavilla et al. (2022) estimate elasticities in the range of 1 to 2 for corporate loans in European countries using interest rate variation during the COVID-19 crisis; these are closer to our estimates.

At the same time, our estimated elasticity is low enough to allow banks to exert substantial ex ante market power. Under monopoly conditions (i.e., $\theta = 1$ in Section 3), this elasticity corresponds to a markup ratio of 2.43 over the expected cost of credit, net of any future surplus from the lending relationship. Thus, even under moderate levels of competition ($0 < \theta < 1$), banks may retain significant pricing power, enabling them to earn high profit margins and reduce equilibrium credit provision relative to the efficient benchmark.

Placebo Analysis Using (Noncapped) Individual Loans. One potential concern with our empirical design is that bank branches with higher exposure to the cap may have been on differential trends in credit market outcomes, independent of the policy. While the absence of pretrends in interest rates and corporate credit provision alleviates this concern, it remains possible that branches with greater treatment intensity experienced abrupt, idiosyncratic shifts in credit conditions around the time of the policy, which may confound our estimates.

To address this issue, we conduct a placebo analysis using individual loans. The core idea is that, since the interest rate cap applied only to corporate loans and not to individual borrowing, we should not observe any systematic relationship between treatment intensity and changes in interest rates or loan volumes for individual borrowers after the introduction of the cap. If such a relationship were present, it would suggest that treatment intensity may be proxying for other unobserved, branch-specific factors affecting credit provision.

Figure 6 presents results from reestimating Equation (5), using interest rates and lending to individuals —comprising loans to sole proprietors and entrepreneurs, as well as consumer installment loans—as the outcome variables. We find no statistically significant effects on interest rates (Panel A) or on the volume of individual credit (Panel B). These null effects reinforce our interpretation that the observed increase in corporate lending was not driven by broader branch-specific trends in credit market outcomes but instead reflects a causal response to the interest rate cap.[17]

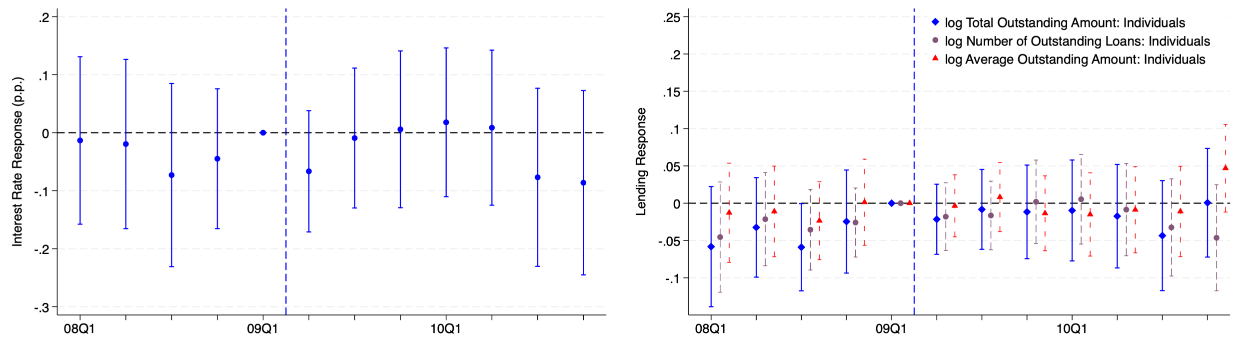
Robustness Analysis. In Appendix B, we provide a battery of robustness checks for our results on the positive effects on credit provision, which we summarize here.

Appendix Figures B.1 and B.2 demonstrate that our results remain robust when controlling

[17]The null effect on individual loan interest rates and lending volumes also alleviates the concern that banks relabeled corporate loans as individual loans and keep charging interest rates above the statutory cap. It also suggests that lenders did not engage in cross-loan product pricing in response to the cap.

Figure 6: Placebo Test: Event-Study Effects of Cap on Loans to Individuals

Panel A: Interest Rate for Individual Loans **Panel B: Credit Provision for Individual Loans**



Notes: This figure plots the estimated coefficients of treatment intensity interacted with quarter dummies from event-study regression (5) with measures of lending to individual borrowers as the outcome. The event-study analysis forms a placebo test, since there was no interest rate cap placed on noncorporate loans. We consider three measures of equilibrium branch-level credit provision: the log of total outstanding loan dollars, the log number of outstanding loans (extensive margin), and the log average outstanding amount (intensive margin), computed as the total outstanding loan amount divided by the number of loans with positive balances remaining. We omit the quarter before the cap reform (2009Q1) as the reference category. Bangladesh Bank announced the interest rate cap on April 19, 2009 (2009Q2), with the cap effective immediately. 95% confidence interval bars are obtained from clustering standard errors at the branch level.

for pre-period trends in interest rates and including subdistrict-by-quarter fixed effects, which control for additional unobserved trends at the branch level.

In Appendix Figure B.3, we present results from a version of our event-study regression (5) where we discretize treatment intensity into quartile bins. We find that the effects on interest rates and credit provision monotonically increase in the treatment intensity. Therefore, our results are not driven by outlier samples in treatment intensity.

In Appendix Figure B.4, we conduct Rambachan and Roth (2023) tests in which we vary the parameter M, representing the multiple by which the posttreatment violations of parallel trends can deviate from the pretreatment differences in trends. The estimated first-stage effect on interest rates remains statistically significant even if we impose a relatively conservative value of $M = 2$. The positive branch-level outstanding loan response remains significant at the 90% level for $M = 2$ and significant at the 95% level for $M < 2$.

In Appendix Figure B.5, we extend the event-study specification in Equation (5) through 2011Q4, covering the period after the partial removal of the interest rate cap (Section 2.1). Because we lack detailed information on which specific loan categories were affected by the partial lifting, we use the same definition of corporate loans as in our baseline analysis. We find that both the decline in interest rates and the increase in credit provision plateaued following the cap's partial removal, consistent with the interpretation that the regulatory change had offsetting effects. Further, given that the rate cap was set at a time when the central bank policy rate was relatively elevated, it is likely that the cap became less binding for new contracts by the time it was rescinded. By 2011Q1, the policy repo (repurchase) rate had fallen from its precap rate level of 8.75% to 5.5%.

Finally, Appendix Table B.1 shows that our results are robust to excluding Islamic banks, which operate under distinct principles prohibiting certain financial transactions and instead apply alternative forms of credit and deposit fees (Choudhary and Limodio 2022).

4.4 Heterogeneous Effects on Credit Provision

While our findings so far indicate that the credit provision increased on average, the cap may have led to credit rationing at some branches. For example, branches facing more competition may have downsized their lending operations, while other branches expanded their credit provision, such thatthe net effect is positive. In this section, we explore heterogeneous effects with respect to branch, subdistrict, and bank characteristics.

In particular, we estimate a modified version of our IV regression (6) to assess heterogeneous semielasticities with respect to interest rates:

$$\begin{aligned} Y_{i,t} =& \alpha \cdot InterestRate_{i,t} + \varphi' \cdot \mathbf{Z_i} \times InterestRate_{i,t} + \sum_{s=-m, s\neq -1}^{s=n} \gamma_s' \cdot \mathbf{Z_i} \times \mathbb{1}\{t=s\} \\ & + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \epsilon_{i,t} \end{aligned} \tag{7}$$

where $\mathbf{Z_i}$ denotes a vector of branch-level characteristics, including attributes of the subdistrict or bank to which branch i belongs. This specification is identical to regression (6) except for the second and third terms. The second term captures heterogeneity in the semielasticity of outcomes with respect to interest rates. The third term controls for time-varying effects of the characteristics $\mathbf{Z_i}$, ensuring that the coefficients φ are identified solely from the quasi-experimental variation induced by the cap. Extending the previous specification, we use $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$ as an IV, and its interaction with $\mathbf{Z_i}$, $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\} \times \mathbf{Z_i}$, as additional instruments. We compute the Lewis and Mertens (2022) first-stage F-statistics, which are autocorrelation- and heteroskedasticity-robust even in the presence of multiple endogenous regressors.

Table 4 presents our results. In column (1), we examine heterogeneity with respect to subdistrict-bank HHI, defined analogously to Section 2.4. The coefficient on the interaction between interest rates and the above-median HHI indicator is statistically insignificant, and its magnitude is modest relative to the main effect. These findings suggest no systematic relationship between the IV effects and local market concentration of the bank.[18]

At first glance, our finding of null heterogeneous effects with respect to HHI may appear contradictory to our finding that the HHI is positively associated with the level of precap interest rates in Section 2.4. However, these findings are, in fact, perfectly compatible with each other. While banks' market power may affect the level of markups, and hence the level of precap rates, it may not be related one-to-one with the slope of credit demand, which is what the IV estimates capture under the assumption that the cap is above break-even rates. Moreover, existing research highlights that HHI is, at best, an imperfect proxy for market power (De Loecker et al. 2020). In Section 5, we provide additional evidence of limited competition by examining the cap's effects on local competing branches, even after controlling for own-branch exposure.

[18] Appendix Figure C.1 explores heterogeneity in the event-study reduced-form estimates across quartiles of lending HHI. We observe no clear pattern of treatment-effect heterogeneity with respect to HHI.

Table 4: Heterogeneous IV Effects of Cap-Induced Changes on Credit Provision

	log Total Outstanding Amount				
	(1)	(2)	(3)	(4)	(5)
Interest Rate	-0.31***	-0.20***	-0.33***	-0.39***	-0.34**
	(0.06)	(0.06)	(0.09)	(0.09)	(0.14)
Interest Rate X Above Median Dummy	0.09				0.04
(Subdistrict-Bank HHI)	(0.08)				(0.09)
Interest Rate X Above Median Dummy		-0.21**			-0.15
(Population Density)		(0.08)			(0.10)
Interest Rate X Above Median Dummy			0.10		0.13*
(Delinquency Rate: $\geq$ 9 Months)			(0.08)		(0.07)
Interest Rate X Above Median Dummy			0.02		0.04
(Proportion of Secured Loans)			(0.07)		(0.08)
Interest Rate X Above Median Dummy			-0.03		-0.09
(Deposit Rate: Total Accounts)			(0.11)		(0.13)
Interest Rate X Above Median Dummy				0.01	0.10
(Leverage Ratio)				(0.12)	(0.12)
Interest Rate X Above Median Dummy				0.13	0.12
(Bank Delinquency Rate: $\geq$ 9 Months)				(0.11)	(0.14)
Specification	IV	IV	IV	IV	IV
Lewis and Mertens F-Statistics	74.50	71.04	32.22	26.18	9.01
Branch FE	X	X	X	X	X
Bank X Quarter FE	X	X	X	X	X
District X Quarter FE	X	X	X	X	X
Baseline Mean	6.91	6.91	6.91	6.91	6.91
Number of Banks	39	39	39	39	39
Number of Branches	1855	1855	1855	1855	1855
Observations	22260	22260	22260	22260	22260

Notes: This table reports results from estimating the IV regression (6) via TSLS with additional terms interacting the endogenous variable, the branch-level average interest rate, with precap branch-level characteristics. We instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch *i* is exposed to the interest rate cap regulation. Likewise, we instrument each interaction term with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$ interacted with the branch-characteristic variable. We define population density and lending HHI as in Table 1. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. We define the leverage ratio as the 2008 year-end ratio of total debt outstanding to total assets at the parent-bank level. The final row refers to the interaction term between interest rates and the nine-month delinquency rate at the parent-bank level. Robust standard errors clustered at the branch level are in parentheses. For each specification, we report the Lewis and Mertens (2022) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instruments. *p<0.1; **p<0.05; ***p<0.01.

In column (2), we investigate heterogeneity in the IV estimates with respect to population density. We find a significantly negative coefficient of -0.21 on the interaction term between interest rates and the dummy for the above-median population density. The magnitude is comparable with the base effect of -0.20, which is also statistically significant. Therefore, the semielasticity of credit demand is approximately twice as large in more populous (urban) areas. This finding is consistent with an interpretation that there are more investment opportunities on the margin in those areas. This finding aligns with our earlier result that the precap interest rates were lower in more populous areas, even conditional on HHI (Table 1). A higher demand elasticity implies that banks charge lower markups, and hence lower interest rates, in more populous areas.[19]

In column (3), we examine heterogeneous IV effects with respect to measures of bank lending risk and deposit rates. We consider interaction terms with delinquency rates, the share of secured loans, and deposit rates, all measured in the precap period. None of these interaction terms is statistically significant, and their magnitudes are relatively modest. These results suggest that credit rationing did not occur even among riskier branches or branches with higher funding costs, as proxied by deposit rates.

In column (4), we explore the heterogeneous IV effects with respect to banks' financial solvency, using precap measures at the parent-bank level: the leverage ratio and the delinquency rate. We adopt the leverage ratio, defined as total bank debt divided by the book value of assets, as our measure of a bank's financial distress (Mayes and Stremmel 2012). Again, we find no statistically significant interaction effects, and the magnitudes are modest.

In column (5), we jointly assess heterogeneous IV effects with respect to all branch-level characteristics considered in columns (1)–(4). Although the first-stage F-statistics are somewhat low in this specification, the main conclusion remains consistent. We continue to find no

[19] Appendix Table C.1 shows that this heterogeneous effect is driven more strongly by the intensive margin (average outstanding loan per account) than the extensive margin (number of loans).

systematic evidence of credit rationing along observable dimensions of branch characteristics.

4.5 Effects on Other Credit Market Outcomes

In this section, we assess the effects of the cap on additional credit market outcomes.

Risk Composition of Borrowers. One common criticism of interest rate caps is that they may lead to rationing of credit for ex ante riskier borrowers. If banks are unable to charge higher interest rates to riskier clients, they may instead prioritize lending to safer, inframarginal borrowers—those who would have received interest rates below 13% even in the absence of the cap. While we find no heterogeneous effects in credit provision with respect to a branch's risk profile (Table 4), it remains possible that branches more exposed to the cap engaged in systematic risk-shifting.

In Table 5, we directly assess this possibility. Column (1) reports estimates from IV regression (6) using the delinquency rate, defined as the share of loans overdue by more than nine months, as the outcome variable. We estimate an effect of 0.002, which is statistically insignificant and small relative to the baseline mean of 0.044. Thus, we find no evidence that the cap led to higher delinquency rates, supporting the interpretation that there was no increased rationing of risky borrowers. In column (2), we repeat the analysis using a less stringent definition of late repayment—loans overdue by more than three months. Again, we find no significant change in the delinquency rate.

Column (3) examines an alternative proxy for loan risk: the fraction of loans backed by collateral. If banks are more likely to require collateral for riskier borrowers, and if the cap reduced the share of such loans, we might observe a decline in the proportion of secured loans. Alternatively, banks may have adjusted collateral requirements in response to lower interest rates to anticipate changes in underlying default risk. We find no significant effect; the point estimate is 0.006, statistically insignificant and small relative to the baseline mean of 0.641.

Table 5: IV Effects of Cap-Induced Changes in Interest Rates on Delinquency and Collateralization

	Delinquency Rate: $\geq$ 9 Months	Delinquency Rate: $\geq$ 3 Months	Proportion of Secured Loans	Predicted Delinquency Rate
	(1)	(2)	(3)	(4)
Interest Rate	0.002	-0.002	0.006	-0.000
	(0.004)	(0.008)	(0.020)	(0.001)
Specification	IV	IV	IV	IV
Montiel Olea and Pflueger F-Statistics	109.538	109.538	109.538	109.538
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	0.044	0.056	0.641	0.057
Number of Banks	39	39	39	39
Number of Branches	1855	1855	1855	1855
Observations	22260	22260	22260	22260

Notes: This table reports results from estimating the IV regression (6) via TSLS for an expanded set of outcome variables, including the 9-month and 3-month delinquency rates, the fraction of loans secured by physical collateral, and the predicted delinquency rate. The predicted delinquency rate is defined as the weighted average of bank-sector delinquency rates, using the outstanding loan amounts by sector as weights. We instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

In column (4), we use a yet another risk measure based on sectoral composition of loans. We construct the average precap delinquency rate for each of 46 detailed sectors at the bank level. We then compute a branch-quarter-level predicted delinquency rate as the weighted average of these bank-sector delinquency rates, using the outstanding loan amounts by sector as weights. This proxy captures whether banks shifted lending toward riskier sectors. The results show no meaningful effect; the estimate is close to zero and precisely estimated relative to the baseline mean.

Taken together, these findings further support the conclusion that the interest rate cap did not lead to credit rationing among riskier borrowers, nor did it induce a shift in borrower risk composition at affected branches.

Table 6: IV Estimates of Cap-Induced Changes in Interest Rates on the Deposit Market

	Deposit Rate: Total Accounts	log Deposit Amount: Total Accounts	Deposit Rate: Corporate Accounts	log Deposit Amount: Corporate Accounts
	(1)	(2)	(3)	(4)
Interest Rate	0.01	-0.04	0.01	0.19
	(0.01)	(0.03)	(0.04)	(0.16)
Specification	IV	IV	IV	IV
Montiel Olea and Pflueger F-Statistics	107.25	107.25	46.96	46.96
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	5.910	6.491	5.934	4.660
Number of Banks	39	39	39	39
Number of Branches	1855	1855	1712	1712
Observations	22260	22260	13035	13035

Notes: This table reports results from estimating the IV regression (6) via TSLS, where now the set of outcome variables includes deposit rates and deposit amounts across all types of accounts (corporate plus individual) and just corporate accounts. As before, we instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

Deposit Market. We also examine the effects of the cap on the deposit market. If banks possess market power not only in lending but also in deposit-taking, as observed in the literature on the deposit franchise in the United States (Drechsler et al. 2017), they may respond to the lending rate cap by adjusting deposit rates downward. However, in Table 6, we find no statistically significant changes in deposit rates or amounts, either at the overall branch level (columns 1 and 2) or specifically among corporate depositors (columns 3 and 4). Combined with the absence of significant changes in risks as mentioned above, our results suggest that bank branches experienced reduced profit margins due to the interest rate cap. Further, given that our placebo analysis (Figure 6) uncovers no changes in credit provision for loans to individuals, which were not subject to the cap, our results point to limited cross-product price spillovers at the branch level. This is an important consideration given concerns about spillovers to unregulated product markets for government loan subsidies when banks wield market power (Haas Ornelas et al. 2024).

Table 7: IV Estimates of Cap-Induced Changes in Interest Rates on Sectoral Composition

	Share of Outstanding Loans: Manufacturing	Share of Outstanding Loans: Import or Export	Share of Outstanding Loans: Others
	(1)	(2)	(3)
Interest Rate	-0.01	-0.00	0.02
	(0.01)	(0.01)	(0.01)
Specification	IV	IV	IV
Montiel Olea and Pflueger F-Statistics	109.54	109.54	109.54
Branch FE	X	X	X
Bank X Quarter FE	X	X	X
District X Quarter FE	X	X	X
Baseline Mean	0.16	0.09	0.75
Number of Banks	39	39	39
Number of Branches	1855	1855	1855
Observations	22260	22260	22260

Notes: This table reports results from estimating the IV regression (6) via TSLS, where now each outcome is the share of outstanding loans for one of the three broad corporate sector classifications: manufacturing, import or export firms, and other sectors. As before, we instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

Sectoral Composition of Borrowers. We now investigate an alternative dimension through which the cap may change banks' allocation of credit across borrower types. In response to the cap, banks may have systematically changed how much they lent to certain sectors depending on their profitability, default probability, or ease of monitoring. In Table 7, we present results from estimating IV regression (6), using the share of outstanding loans toward three broad categories of borrower sectors: manufacturing, import/export, and other. Mechanically, the regression coefficients sum up to zero. We find no evidence that the branch-level average shares of loans disbursed to any of the three major sectors receiving loans regulated by the cap changed in response to the cap reform.[20]

[20] The baseline branch-level means for the shares of manufacturing and import/export sectors (16 and 9%) are substantially lower than their corresponding aggregate shares across all branches (46 and 17%; Appendix Figure A.1). This discrepancy arises because lending to these sectors is concentrated in a relatively small number of branches located in urban areas.

5 Additional Evidence of Banks' Ex Ante Market Power

Our results so far are consistent with the presence of ex ante market power in the corporate banking sector. In this section, we provide further complementary evidence that banks in Bangladesh exercise their ex ante market power by charging markups over the break-even interest rate on loans.

5.1 Effects of Cap-Induced Changes in Competitors' Interest Rates

In our first empirical exercise of this section, we show that changes in interest rates among branches attached to close-competitor banks operating in the same local market do not affect credit market outcomes. If banks are subject to fierce competition for borrowers, a reduction in interest rates by a competing bank branch should lower own-branch credit demand. However, if borrowers cannot flexibly substitute across lenders, each bank branch operates as a monopoly, and thus the reduction in the competing bank branch's interest rates does not affect own-branch credit demand.

An obvious concern for implementing this test using OLS is that a competing bank branch's interest rate is likely to be set endogenously in response to local credit demand. To deal with this concern, we use the interest rate cap policy to induce variation in competing bank branch interest rates. Specifically, we study how cap-induced changes in loan rates charged by competing bank branches in the same subdistrict affect credit demand, after controlling for the branch's own cap-induced interest rate changes. We implement this design by augmenting our IV regression (6) to include a separate term, $CompetingRate_{i,t}$, measuring the average interest rate charged on regulated loans across any local close-competitor branches to branch i:

$$Y_{i,t} = \alpha_1 \cdot InterestRate_{i,t} + \alpha_2 \cdot CompetingRate_{i,t} + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \epsilon_{i,d,t} \quad (8)$$

To construct $CompetingRate_{i,t}$, we begin by identifying the closest competing bank for the parent bank of branch i. More concretely, we compute the Euclidean distance between each pair of banks based on two sets of characteristics from the precap period: aggregate balance-sheet size and the vector of outstanding loan shares across sectors. This distance metric allows us to identify which bank is most similar in terms of size and lending specialization. Once the closest competitor bank is identified, we locate any branch of the competitor bank operating in the same subdistrict as branch i and record the interest rates charged by that branch. For this analysis, we restrict the sample to branches having a consistent local competitor branch over the full sample period, such that $CompetingRate_{i,t}$ is always well-defined. In instances where branch i's competing bank has multiple branches in the same subdistrict (550 of 1,084 branches with a nearby competitor branch), we take a share-weighted average of interest rates charged across the local competing branches. By focusing on competitor banks that are similar in both scale and sectoral orientation and restricting the comparison geographically, this procedure aims to capture the most relevant alternative for borrowers at branch i.

Following our baseline IV regression (6), we instrument $InterestRate_{i,t}$ using a branch's exposure to the interest rate cap, defined as $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$. Analogously, we instrument $CompetingRate_{i,t}$ with $TrtIntensity_j \times \mathbb{1}\{t \geq 2009Q2\}$, where j denotes the closest competitor's branch or set of competing branches identified through the procedure described above. The coefficient α_2 thus captures the additional effects of cap-induced changes in competitor interest rates, over and above the effects stemming from the branch's own rates.

Table 8 reports the results from the IV regression (8), using credit provision as the outcome variable. Consistent with our main IV results in Table 3, we find negative and statistically significant coefficients on the branch's own interest rate. In contrast, the coefficient on the competing branch's interest rate is close to zero and statistically insignificant. Notably, the standard errors for both coefficients are of similar magnitude, suggesting that the null result for the competing rate is not due to imprecise estimation.

Table 8: Null IV Effects of Cap-Induced Rate Changes by Competitor Branch

	log Total Outstanding Amount	log Number of Outstanding Loans	log Average Outstanding Amount
	(1)	(2)	(3)
Interest Rate	-0.34***	-0.13*	-0.22***
	(0.08)	(0.07)	(0.06)
Competing Branch's Interest Rate	-0.12	-0.11	-0.01
	(0.07)	(0.06)	(0.06)
Specification	IV	IV	IV
Lewis and Mertens F-Statistics	41.22	41.22	41.22
Branch FE	X	X	X
Bank X Quarter FE	X	X	X
District X Quarter FE	X	X	X
Baseline Mean	7.33	4.40	2.94
Number of Banks	39	39	39
Number of Branches	1084	1084	1084
Observations	13008	13008	13008

Notes: This table reports results from estimating the IV regression (8) via TSLS. We instrument own-branch average interest rates $InterestRate_{i,t}$, with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. We instrument $CompetingRate_{i,t}$ with $TrtIntensity_j \times \mathbb{1}\{t \geq 2009Q2\}$, where j denotes the closest competitor's branch or set of competing branches located within the same subdistrict. We restrict the sample to branches having a consistent local competitor branch over the full sample period, such that $CompetingRate_{i,t}$ is always well-defined. All specifications include bank-quarter and branch fixed effects. Robust standard errors clustered at the branch level are in parentheses. For each specification, we report the Lewis and Mertens (2022) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instruments. *p<0.1; **p<0.05; ***p<0.01.

This finding is consistent with the interpretation that banks have ex ante market power. If banks face fierce local competition for borrowers, changes in interest rates among competing branches should induce borrowers to substitute toward branches offering lower interest rates. On the other hand, if branches do not face strong competition, credit demand may not respond to competing-branch interest rates.

5.2 Effects of Competitor-Bank-Branch Entry on Lending

In our second empirical exercise, we present further evidence on banks' ex ante market power by examining how the entry of a competing bank branch affects the credit demand faced by incumbent branches. The underlying idea is similar to the previous analysis: if branches operate under intense competition, the entry of a nearby competitor should lead to a reduction in credit demand at the incumbent branch due to borrower substitution, which may also influence its interest-rate-setting strategy. In contrast, if banks possess substantial ex ante market power, such entry may have little to no effect on incumbent credit demand.

To test this hypothesis, we estimate the following regression:

$$Y_{i,t} = \sum_{t=-m, t\neq -1}^{t=n} \xi_t \cdot \text{CompetitorEntry}_{i,t} + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \varepsilon_{i,t} \tag{9}$$

where $\text{CompetitorEntry}_{i,t}$ is a dummy equal to unity if branch i in quarter t experienced entry into the same district of a new branch belonging to its parent bank's closest competitor bank, where we define the competitor bank following the same definition as in Section 5.1.[21] Similarly to our regression specification for the effects of the interest rates (5), we control for branch fixed effects η_i, bank-quarter fixed effects $\nu_{Bank(i),t}$, and district-quarter fixed effects $\psi_{District(i),t}$, such that we identify the effects of a competing branch's entry using within-bank and -district variation. We estimate this regression using the postcap sample spanning the 2011Q1–2014Q2 period, such that this result is not confounded by the effects of the rate cap policy that we have studied so far. During this period, 17% of bank branches experienced competing-branch entry. We stack up entry events within each branch, meaning that $\text{CompetitorEntry}_{i,t}$ assigns entry timing according to the first instance of the competitor bank's entry into the subdistrict where i is located.

An important identification concern is that branch entry is unlikely to be random; banks are more likely to open branches in areas where credit demand is strong or growing. While our DiD design accounts for differences in the level of credit demand across locations, the identification strategy may be compromised if banks systematically enter subdistricts experiencing faster

[21] If a branch experienced multiple competitor entry during our sample period, we define the quarter of competitor entry as the first quarter that such event happened. This happened only for a few cases: just 62 of the 2,723 branches in our sample experienced multiple competitor entry events during the sample period.

growth in credit demand, thereby violating the parallel-trends assumption.

To address this concern, we implement the estimator proposed by de Chaisemartin and D'Haultfœuille (2020), which identifies treatment effects by comparing groups treated at different times. The identification assumption is that, conditional on eventual entry, the exact timing of branch entry is as good as random. By explicitly excluding branches that never faced competitor entry during our sample periods, which are predominantly concentrated in urban areas (Figure 3), we ensure that our results are not driven by a simple comparison between urban and rural markets. Furthermore, to alleviate the concern that the timing of competitor entry may be systematically related to unobserved changes in the local credit market, we assess pretrends in outcome variables.[22]

Figure 7 presents our event-study results for interest rates and credit provision. Consistent with the identifying assumption, we observe limited pretrends, suggesting that incumbent branches and soon-to-be-treated branches follow common trajectories prior to entry. Following the entry of a competing branch, we find no statistically significant effects on either interest rates or credit provision. This pattern accords with the interpretation that the entering branch exerts only limited competitive pressure on the incumbent. In Appendix Figure E.2, we conduct a related test in which show that entry of *any* new branch into a subdistrict has no discernible effect on total lending by incumbent branches but leads to an overall increase in total lending in the subdistrict.

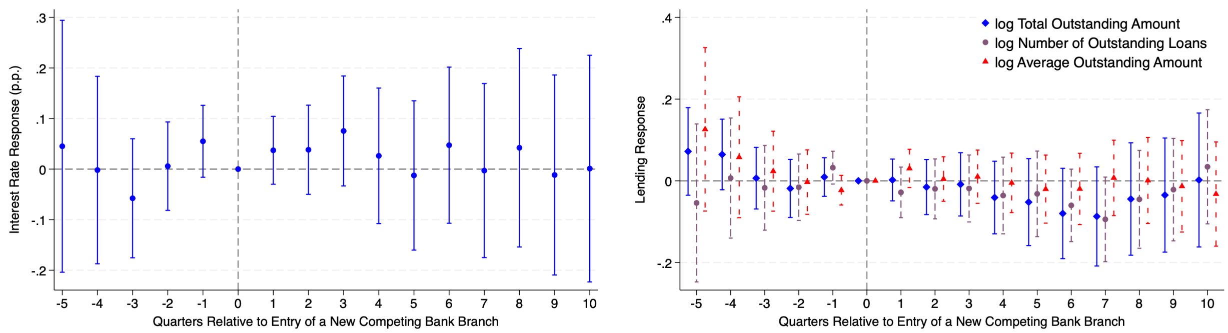
6 Discussion and Policy Implications

Our results are broadly consistent with the interpretation that banks possess substantial ex ante market power. In Section 4, we show that branches more exposed to the interest rate cap

[22]While our preferred specification follows de Chaisemartin and D'Haultfœuille (2020), which relies only on not-yet-treated units as the control group, Appendix Figure E.1 shows that our results are robust to alternative estimators for staggered event-study designs proposed by Callaway and Sant'Anna (2021), Sun and Abraham (2021), and Borusyak et al. (2024), which use never-treated units as additional control groups.

Figure 7: Null Effects of the Entry of a Close Competitor Bank's Branch

Panel A: Interest Rates **Panel B: Credit Provision**



Notes: This figure plots the estimated quarterly event-study coefficients on the sequence of dummies $\text{CompetitorEntry}_{i,t}$ from regression (9) via the de Chaisemartin and D'Haultfœuille (2020) estimator, with 1,000 bootstrap iterations to compute standard errors.. We consider the four main outcome variables we use throughout this paper: share-weighted average interest rates in Panel A; log outstanding loan amounts, the log number of loans, and the log average loan size in Panel B. In each regression, we omit the quarter prior to the competing branch's entry as the reference category. 95% confidence interval bars are obtained from clustering standard errors at the branch level.

increased their credit provision, without corresponding increases in measured risk or funding costs. Section 5 further documents the absence of credit demand responses to competitors' cap-induced interest rate changes or competitor entry, reinforcing the view that banks hold significant ex ante market power. While ex post market power may be warranted to sustain lending under weak contract enforcement, our findings suggest that the ex ante market power brings a distinct distortion toward underprovision of credit, as discussed in Section 3.

Do our results imply that the interest rate cap improved aggregate output and welfare by eliminating the market-power distortions in Bangladesh? While it is tempting to conclude so, we caution against this interpretation for several reasons.

First, our DiD design compares branches with varying exposure to the cap and does not capture general equilibrium or aggregate effects. For instance, increased investment by firms borrowing from more-exposed branches could depress output prices, shifting credit demand nationwide (Saidi and Streitz 2021). Similarly, banks may offset reduced lending margins by adjusting deposit rates (Dubuis and de Fraisse 2024, Wang 2025). While we find no evidence of such adjustments at the branch level (Section 4.5), our inclusion of bank-by-quarter fixed effects limits our ability to detect such responses at the bank level. Banks could also reallocate credit across branches in response to the cap. Although we find no evidence of within-district reallocation (Appendix Table F.1), we cannot rule out shifts across districts. Overall, our empirical design identifies branch-level effects and is not suited to evaluating aggregate consequences.

Second, our analysis focuses solely on the cap's effects on credit provision. In practice, interest rate caps may interact with other policies, such as monetary policy or bank capital requirements. A key concern is that caps may blunt the transmission of monetary policy if lending rates become binding at the cap (International Monetary Fund 2011). Evaluating this trade-off between reducing banks' market power and maintaining monetary policy effectiveness requires a general equilibrium framework (e.g., Wang et al. 2022).

Third, our study captures only the short-run effects of the cap, which remained fully in place for two years. If maintained over longer periods, such caps may distort other aspects of banking markets, such as branch expansion or new bank entry (Ji et al. 2023, Rysman et al. 2023).

While our results do not permit definitive conclusions about the overall policy effects, they do provide clear evidence that ex ante market power distorts the corporate loan market in Bangladesh. These distortions likely contribute to the persistently high interest rates in the corporate lending sector. The potential need for ex post market power to sustain relationship lending does not justify the inefficiencies created by ex ante market power. Thus, addressing ex ante market power distortions should be a key priority for policymakers. Interest rate cap policy may be one potential, though not exclusive, policy tool to help achieve this objective.

7 Conclusion

We study the effects of interest rate cap regulation on the corporate banking market using credit-registry data and a policy experiment in Bangladesh. In 2009, the Bangladesh central bank introduced an interest rate cap for business term loans of 13%. Using DiD designs comparing bank branches more vs. less exposed to the reform according to the proportion of loans that ex ante featured rates above the cap threshold, we document that the introduction of the cap significantly increased equilibrium credit provision along both the intensive and extensive margins. We do not find any change in borrower risk profiles, as proxied by delinquency rates or whether loans were secured via physical collateral. These findings stand in contrast to prior studies of interest rate caps in high- or upper-middle-income countries and in consumer or small business lending, which typically document *declines* in credit supply, especially for riskier borrowers. Our results suggest that ex ante market power is a key distortion in corporate lending markets in low-income countries and highlight the need for context-specific policy prescriptions, distinct from those applied to more developed financial systems.

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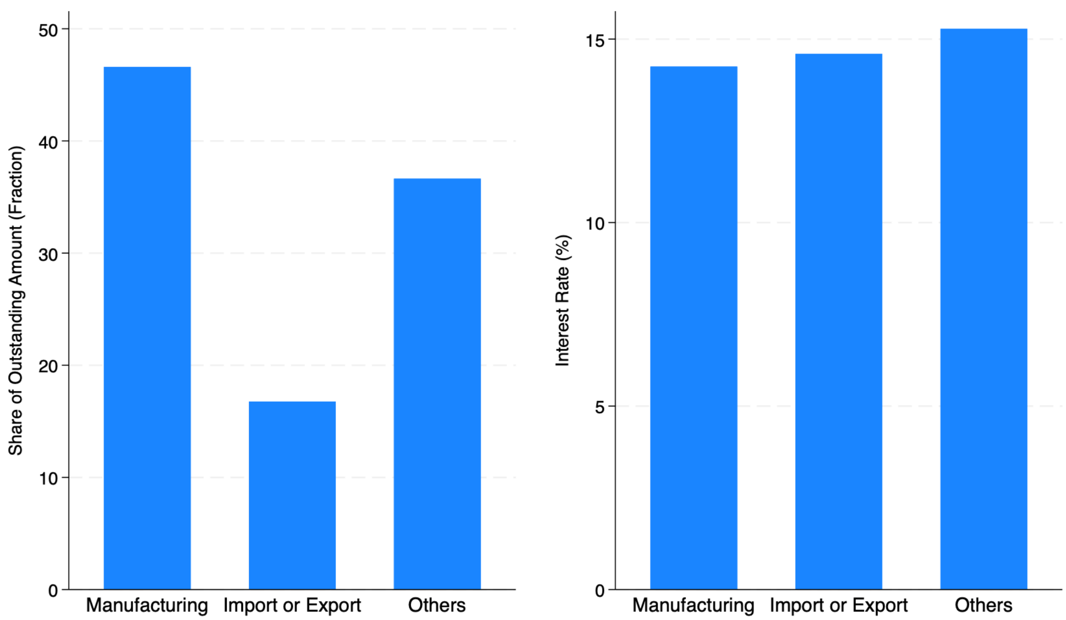
Online Appendix to

Interest Rate Caps, Corporate Lending, and Bank Market Power: Evidence From Bangladesh

by Yusuke Kuroishi (Hitotsubashi University), Cameron LaPoint (Yale SOM), and Yuhei Miyauchi (Boston University)

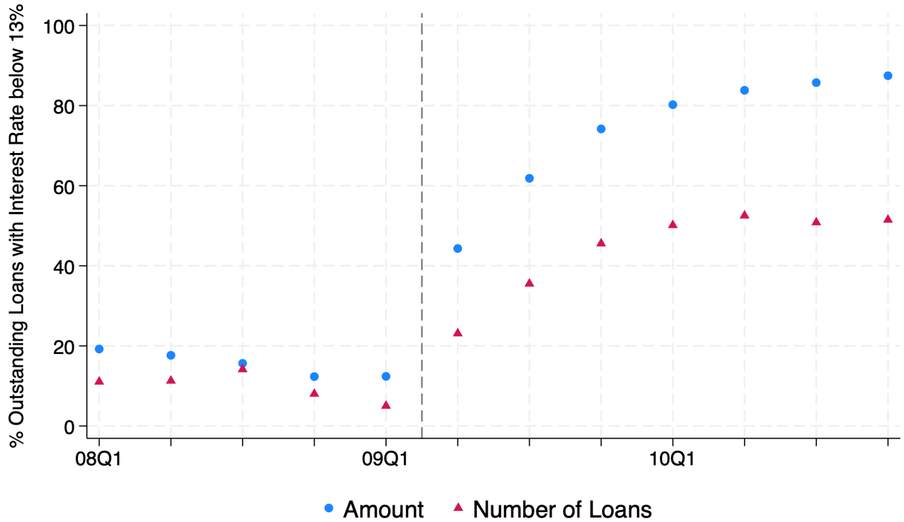
A Appendix for Data and Descriptive Analysis

Figure A.1: Share of Precap Outstanding Loans and Interest Rates by Broad Industrial Sector



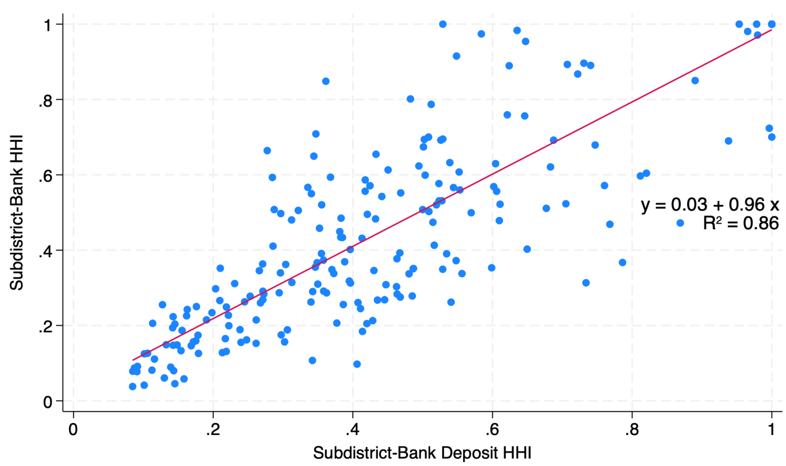
Notes: We tabulate the shares of lending by total outstanding loan amounts (left-hand side) and average annualized interest rates (right-hand side). As in the main analysis, we compute a weighted average of interest rates using outstanding loan amount shares as the weights. We aggregate the 46 sectoral classification codes in the Scheduled Bank Statistics (SBS) data to three broad sectors: manufacturing, import or export firms, and all other firms ("others").

Figure A.2: Proportion of Outstanding Loans with Interest Rates below 13% Cap



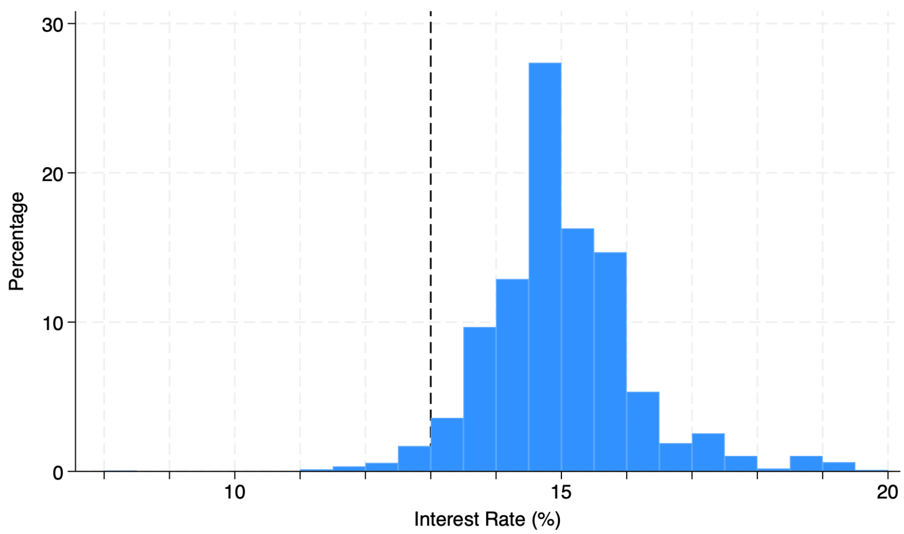
Notes: The figure plots for each quarter the fraction of outstanding loan counts and lending amounts carrying annualized interest rates equal to or below the 13 percent statutory cap. The vertical dashed line indicates the quarter in which the cap was enacted (2009Q2).

Figure A.3: Scatterplot of Subdistrict-Bank Lending and Deposit HHI



Notes: The figure compares market concentration in the lending and deposit markets across branches. Subdistrict-Bank HHI is the Herfindahl-Hirschman Index calculated based on either the market share of the outstanding loan amount lent by a parent bank (y-axis) or the parent bank's share of total savings deposit amounts in a subdistrict prior to the regulation. In calculating lending HHI, we exclude loans with greater than 9 months of delinquent payments. We plot the trend line in red, obtained from regressing the Subdistrict-Bank HHI on the Subdistrict-Bank Deposit HHI.

Figure A.4: Distribution of Precap Interest Rates across Branches



Notes: The figure displays the distribution of branch-level loan share-weighted average annualized interest rates during the precap period. The vertical dashed line indicates the 13 percent statutory cap which was later imposed in 2009Q2.

Table A.1: Patterns of Precap Branch-Level Interest Rates

	Interest Rate				
	(1)	(2)	(3)	(4)	(5)
Deposit Rate	-0.19***	-0.16	-0.15	-0.16	-0.17*
	(0.04)	(0.10)	(0.11)	(0.11)	(0.10)
Delinquency Rate: $\geq$ 9 Months	-1.24***	-0.31*	-0.08	-0.04	-0.09
	(0.30)	(0.17)	(0.16)	(0.16)	(0.17)
Proportion of Secured Loans	-0.81***	0.18***	0.16***	0.16***	0.11*
	(0.08)	(0.06)	(0.06)	(0.06)	(0.07)
ln(Population Density)			-0.12***	-0.10***	
			(0.01)	(0.01)	
Subdistrict-Bank HHI				0.21***	
				(0.06)	
Specification	OLS	OLS	OLS	OLS	OLS
Bank FE		X	X	X	X
Subdistrict FE					X
Mean	14.99	14.99	14.99	14.99	14.99
Number of Banks	39	39	39	39	39
Observations	2124	2124	2124	2124	2124
Adj. R-squared	0.10	0.59	0.62	0.62	0.64

Notes: Each column in the table reports results from estimating a predictive regression of precap branch-level average corporate loan interest rates on branch characteristics. Subdistrict-Bank HHI is the Herfindahl-Hirschman Index calculated based on the market share of the outstanding loan amount lent by a parent bank in a subdistrict prior to the regulation. In calculating HHI, we exclude loans with greater than 9 months of delinquent payments. We compute population density by dividing population by the total land area in a subdistrict.

Table A.2: Patterns of Deposit Rates on Lending and Deposit HHI

	Deposit Rate	
	(1)	(2)
ln(Population Density)	0.000	0.001
	(0.001)	(0.001)
Subdistrict-Bank HHI	0.009	
	(0.010)	
Subdistrict-Bank Deposit HHI		0.018
		(0.011)
Specification	OLS	OLS
Bank FE	X	X
Quarter FE	X	X
Mean	5.901	5.901
Number of Banks	39	39
Number of Branches	2124	2124
Observations	9959	9959
Adj. R-squared	0.885	0.885

Notes: The table extends the analysis in Table 1 to consider how branch-level deposit rates vary with branch characteristics during the precap period. We define Subdistrict-Bank HHI and Subdistrict-Bank Deposit HHI as in the previous figures. Each regression includes bank fixed effects and quarter fixed effects to account for any seasonality in lending.

B Average Effects of Caps on Interest Rates and Credit Supply

In this appendix, we present extensions of our baseline event-study results estimating the reduced form effects of the interest rate cap on interest rates and lending outcomes.

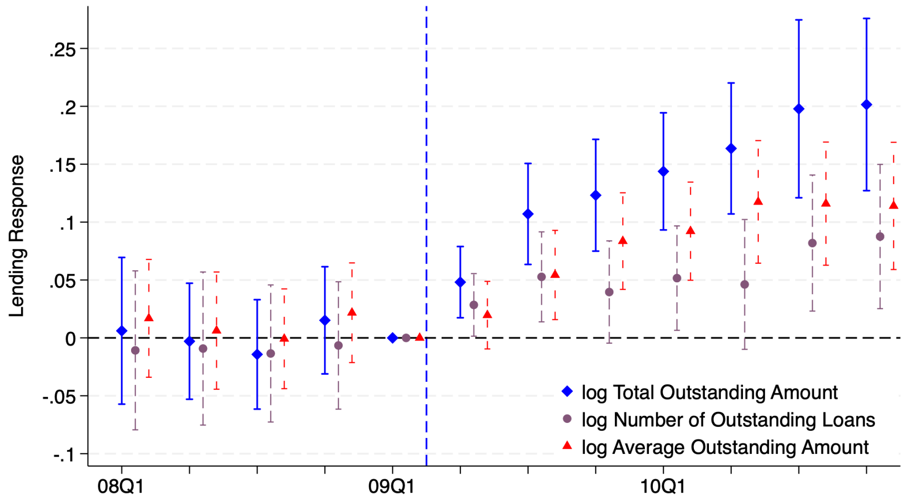
First, Figures B.1 and B.2 demonstrate that our results remain robust when controlling for pre-period trends in interest rates and including subdistrict-by-quarter fixed effects, which control for additional unobserved trends at the branch level.

Second, when we re-estimate the reduced form specification (5) by defining $TrtIntensity_i$ in terms of dummies indicating the quantile of branch-level exposure to the cap, we uncover monotonically increasing effects of the cap on interest rates and lending with respect to the magnitude of treatment (Figure B.3); this indicates that our baseline results are not driven by the linear parameterization of treatment. As described in Section 4.3, we conduct tests for potential violations of parallel trends, following Rambachan and Roth (2023), in Figure B.4. For these tests, we vary the parameter M, representing the multiple by which the posttreatment violations of parallel trends can deviate from the pretreatment differences in trends. The estimated first-stage effect on interest rates remains statistically significant even if we impose a relatively conservative value of $M = 2$. The positive branch-level outstanding loan response remains significant at the 90% level for $M = 2$ and significant at the 95% level for $M < 2$.

Third, we present in Figure B.5 results from our event-study equation (5) estimated over an extended sample period including four quarters after the partial repeal of the cap in 2011Q1. The decline in interest rates and the increase in credit provision plateaued following the cap's partial removal, consistent with the interpretation that the regulatory change had offsetting effects. We conjecture that, since the rate cap was set at a time when the central bank policy rate was relatively elevated, the cap became less binding for new contracts by the time it was rescinded. As of 2011Q1, the policy repo (repurchase) rate had fallen from its precap rate level of 8.75% to 5.5%.

Finally, Table B.1 shows that our results are robust to excluding Islamic banks, which operate under distinct principles prohibiting certain financial transactions and instead apply alternative forms of credit and deposit fees (Choudhary and Limodio 2022).

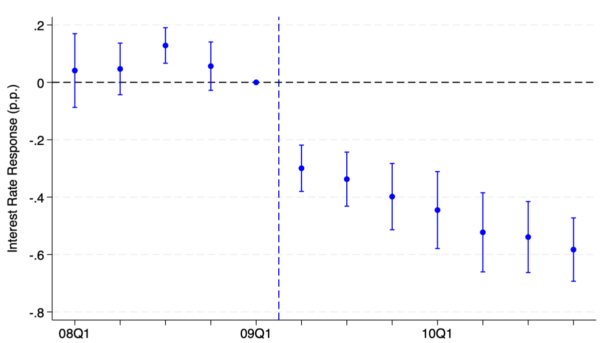
Figure B.1: Event-Study Effects of the Rate Cap on Credit Provision: Controlling for Pre-Period Growth in Branch-Level Interest Rates



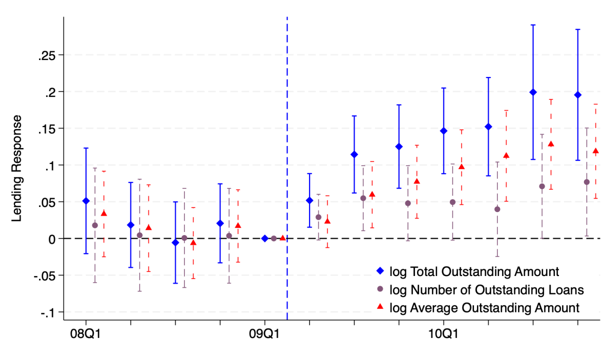
Notes: A version of Figure 5, except that we now augment our main reduced form specification (5) to control for the precap growth of branch-level interest rates (from 2008Q1 to 2009Q1) interacted with quarter dummies. See Section 4.2 of the main text for more details. 95% confidence interval bars obtained from clustering standard errors at the branch level.

Figure B.2: Event-Study Effects of Rate Cap on Interest Rates and Credit Provision: With Subdistrict-Quarter Fixed Effects

(a) **Interest Rates**



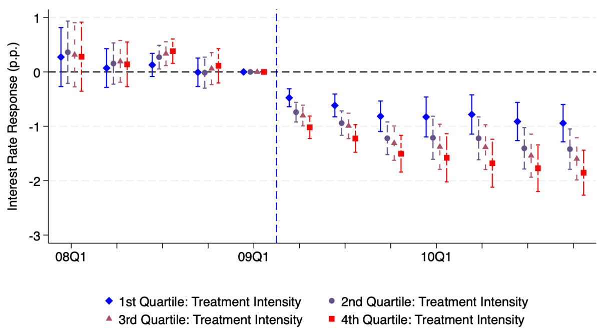
(b) **Credit Provision**



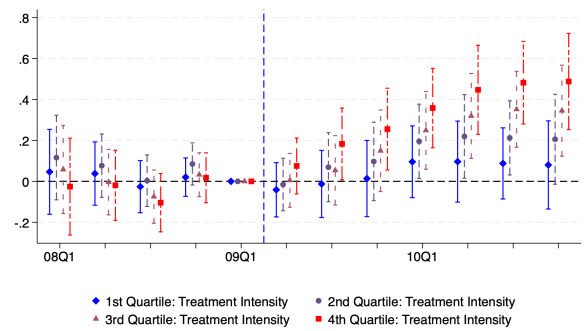
Notes: A version of Figures 4 and 5, except that we replace district-quarter fixed effects with subdistrict-quarter fixed effects to estimate (5). 95% confidence interval bars obtained from clustering standard errors at the branch level.

Figure B.3: Event-Study Effects of Rate Cap on Interest Rates and Credit Provision: By Treatment Intensity Quartile Bins

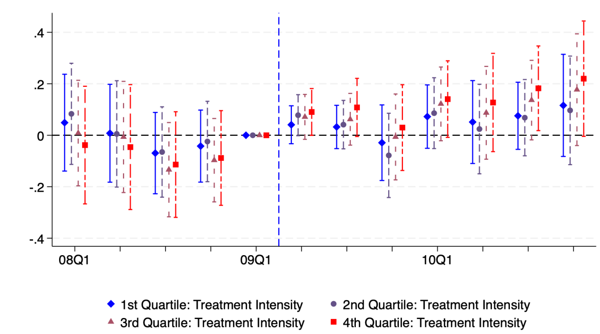
(a) **Interest Rates**



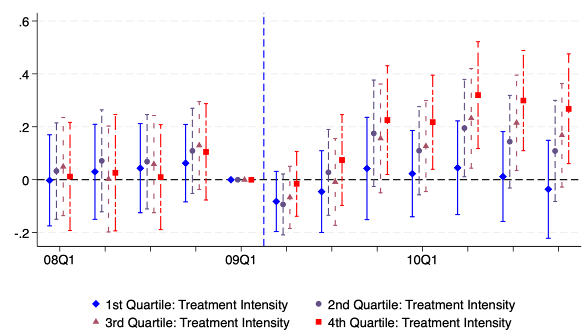
(b) **Log Outstanding Loan Amount**



(c) **Log Number of Loans**



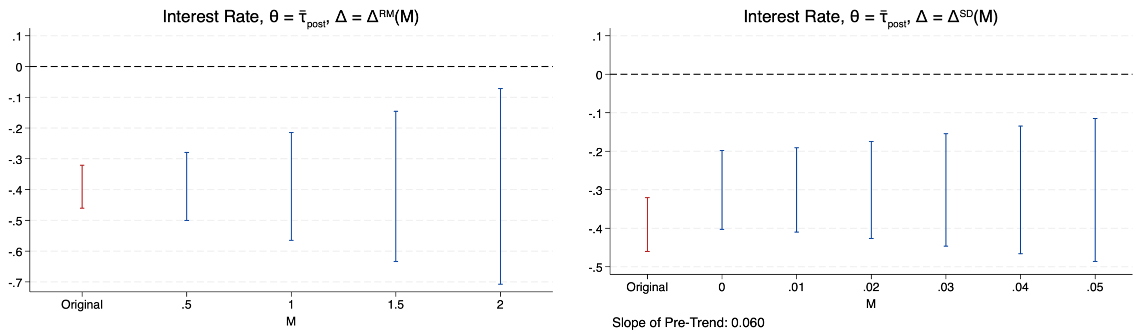
(d) **Log Average Outstanding Loan Amount**



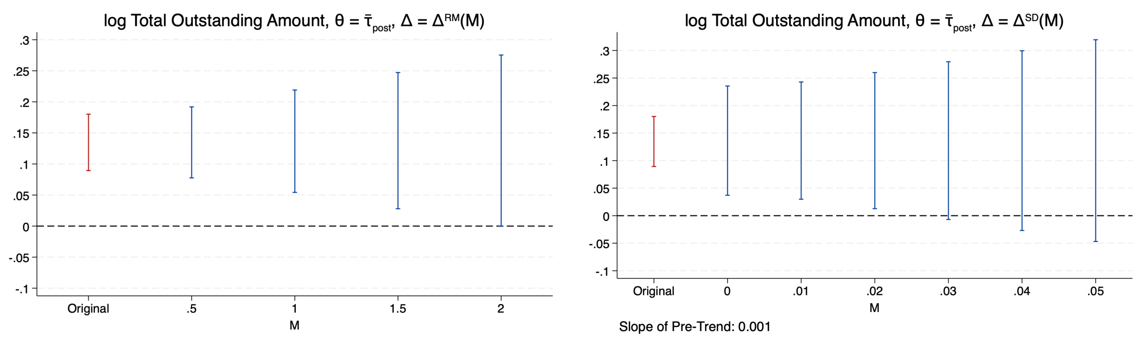
Notes: The figure plots the estimated coefficients of treatment intensity quartile bin dummies interacted with year dummies from the reduced form specification in (5). We continue to omit the quarter directly prior to the reform (2009Q1) as a reference category, as in our main results in Table 2. We consider the four main outcome variables adopted throughout the paper: share-weighted average interest rates in Panel (a); log outstanding loan amounts in (b); the log number of loans in (c); and the log average loan size in (d). 95% confidence interval bars obtained from clustering standard errors at the branch level.

Figure B.4: Rambachan-Roth Robust Pre-Trend Tests for Main Lending Outcomes

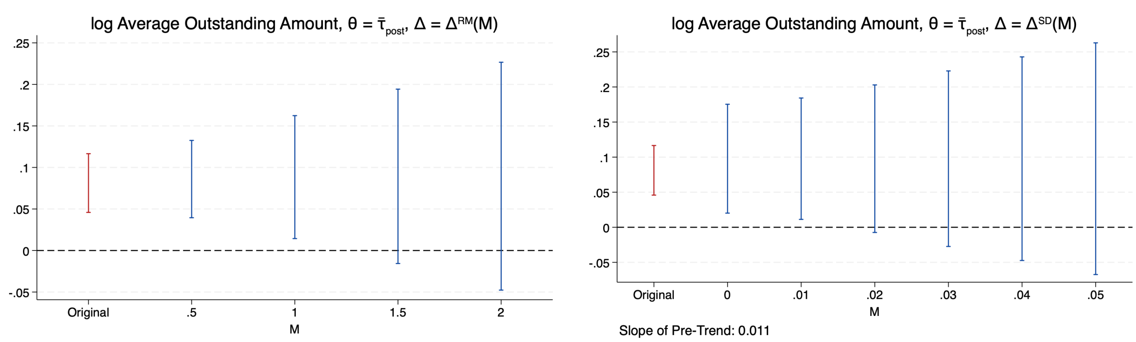
(a) **Interest Rates**



(b) **log Total Outstanding Loans**



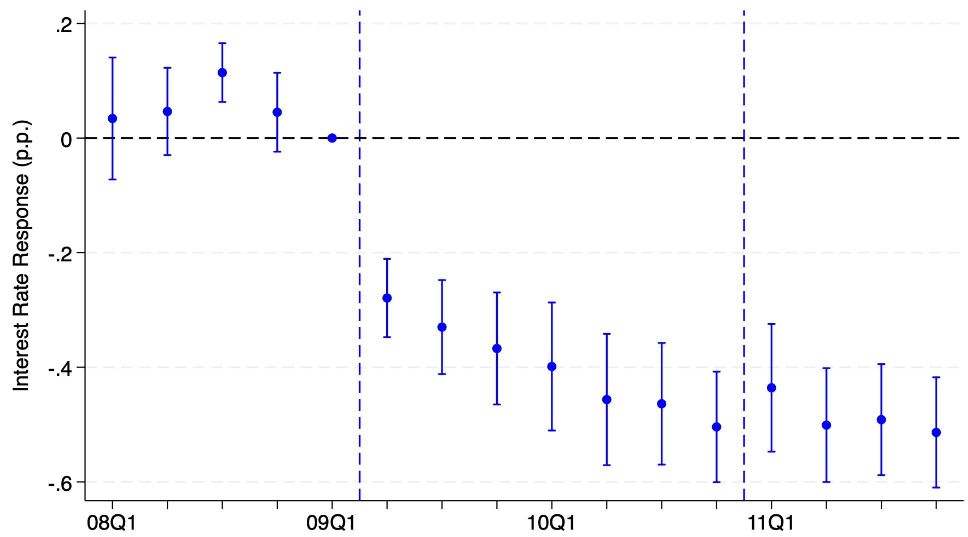
(c) **log Average Outstanding Amount**



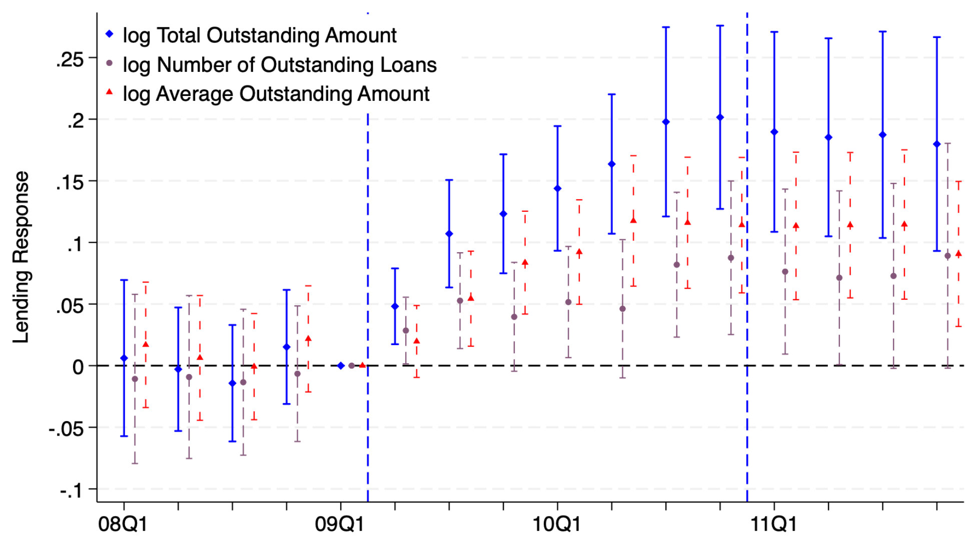
Notes: The figure plots the 95% confidence intervals obtained from the robust pre-trends tests proposed by Rambachan and Roth (2023). We re-estimate the confidence intervals for different values of the parameter M (the x-axis variable), which represents the maximum amount that post-treatment violations of parallel trends can differ from pre-treatment differences in trends such that the treatment effect is partially identified. We report results for the first stage effect of the interest rate cap on branch-level interest rates, and the reduced form effects on log total outstanding loans and log average outstanding loan amounts (intensive margin). We perform pre-trend tests on the estimated dynamic effects pooled over time according to regression equation (5). See corresponding "original" estimates in the odd columns of Table 2 which include bank-quarter and district-quarter fixed effects. Left-hand side panels define M in terms of levels of the point estimate, while right-hand side panels instead define M in terms of the corresponding standard errors.

Figure B.5: Event-Study Effects of Rate Cap on Interest Rates: Extended Period

(a) **Interest Rates**



(b) **Credit Provisions**



Notes: A version of Figures 4 and 5, estimated over the full sample period extending four quarters beyond the partial repeal of the interest rate cap policy in 2011Q1. The figure plots the estimated coefficients of treatment intensity interacted with quarter dummies from event-study equation (5) with the share-weighted average branch-level interest rate charged on corporate loans (panel a) and credit provision (panel b) as the outcome. We omit the quarter before the cap reform (2009Q1) as the reference category. Bangladesh Bank announced the interest rate cap on April 19, 2009 (2009Q2), with the cap effective immediately. 95% confidence interval bars obtained from clustering standard errors at the branch level.

Table B.1: Event-Study Effects of Rate Cap on Interest Rates and Credit Provision: Excluding Islamic Banks)

	Interest Rate	log Total Outstanding Amount	log Number of Outstanding Loans	log Average Outstanding Amount
	(1)	(2)	(3)	(4)
TrtIntensity x 2008Q1-Q4	0.06*	0.01	0.00	0.01
	(0.03)	(0.03)	(0.03)	(0.02)
TrtIntensity x 2009Q2-Q4	-0.32***	0.09***	0.04**	0.05***
	(0.04)	(0.02)	(0.02)	(0.02)
TrtIntensity x 2010Q1-Q4	-0.41***	0.16***	0.06**	0.10***
	(0.05)	(0.03)	(0.03)	(0.03)
Specification	OLS	OLS	OLS	OLS
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	14.88	6.77	4.09	2.68
Number of Banks	31	31	31	31
Number of Branches	1468	1468	1468	1468
Observations	17616	17616	17616	17616
Adj. R-squared	0.87	0.95	0.89	0.91

Notes: We replicate the analysis conducted in Table 2 but now exclude the eight Islamic finance banks from the estimation sample. All regressions in the table include bank-quarter and district-quarter fixed effects. Robust standard errors clustered at the branch level are in parentheses. *p<0.1; **p<0.05; ***p<0.01.

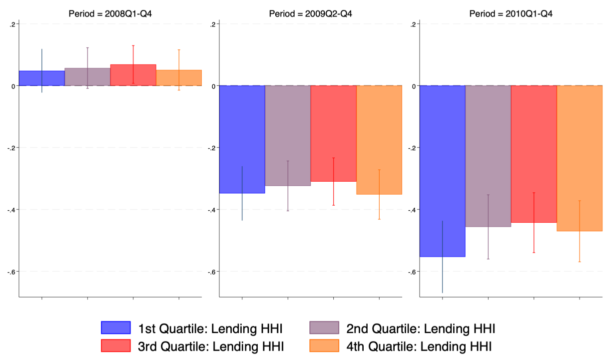
C Heterogeneous Effects of Caps on Interest Rates and Credit Provision

In this appendix, we examine heterogeneous responses of branches to the cap, extending our analysis in Section 4.4 along several dimensions. We show in Figure C.1 that there is no statistically distinguishable heterogeneity in how branch-level interest rates or lending responds to the cap depending on subdistrict-bank lending HHIs. Given the evidence in Figure A.3 that subdistrict-bank lending and deposit HHIs have a nearly 100% correlation in the cross-section of branches, this also implies that there is no meaningful heterogeneity in responses by deposit market concentration.

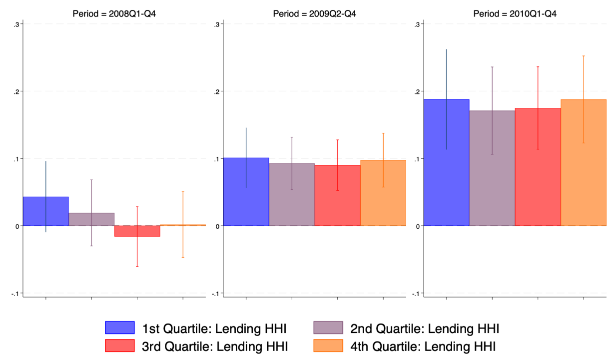
Table C.1 shows that there is little heterogeneity in extensive margin lending responses according to branch-level characteristics. On the intensive margin of lending, the semielasticity of credit demand is over twice as large in more populous (urban) areas. We find marginally statistically significant evidence that financial constraints of banks may play a role in intensive margin lending responses; banks with high leverage ratios do not offer larger loans in response to the reform, compared to their counterparts with low leverage ratios.

Figure C.1: Heterogeneous Reduced Form Responses to the Cap by Lending HHI

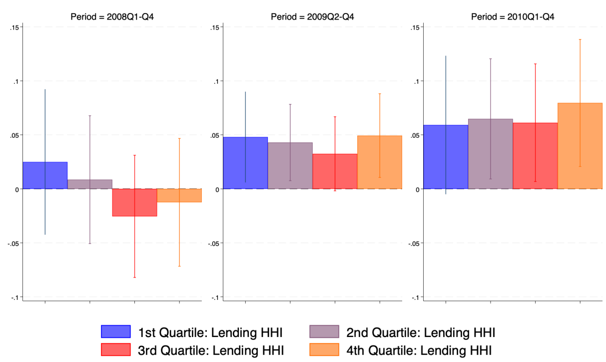
(a) **Interest Rates**



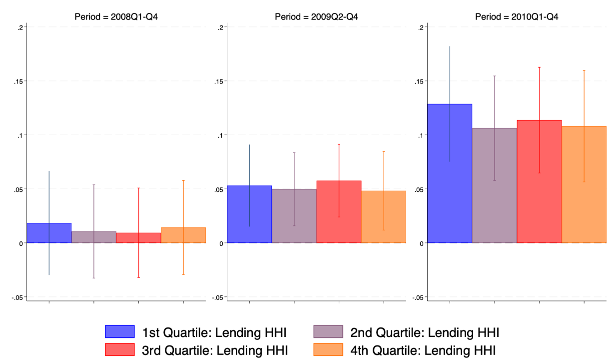
(b) **Log Outstanding Loan Amount**



(c) **Log Number of Loans**



(d) **Log Average Outstanding Loan Amount**



Notes: The figure plots the estimated coefficients of treatment intensity interacted with year dummies from the reduced form specification in (5), for branches within each quartile of lending Herfindahl-Hirschman Index (HHI). We continue to omit the quarter directly prior to the reform (2009Q1) as a reference category, as in our main results in Table 2. We define the lending HHI at the subdistrict-bank (*upazila*) level based on loan dollars outstanding. We consider the four main outcome variables adopted throughout the paper: share-weighted average interest rates in Panel (a); log outstanding loan amounts in (b); the log number of loans in (c); and the log average loan size in (d). 95% confidence interval bars obtained from clustering standard errors at the branch level.

Table C.1: IV Estimates of Heterogeneous Effects on Credit Provision by Branch Characteristics

	log Number of Outstanding Loans					log Average Outstanding Amount				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Interest Rate	-0.11*	-0.11*	-0.14*	-0.09	-0.09	-0.20***	-0.09*	-0.19**	-0.23***	-0.19*
	(0.06)	(0.06)	(0.08)	(0.07)	(0.10)	(0.05)	(0.05)	(0.07)	(0.07)	(0.10)
Interest Rate X Above Median Dummy	0.00				0.02	0.09				0.02
(Subdistrict-Bank HHI)	(0.08)				(0.08)	(0.06)				(0.06)
Interest Rate X Above Median Dummy		-0.07			-0.01		-0.14***			-0.17***
(Population Density)		(0.08)			(0.08)		(0.05)			(0.07)
Interest Rate X Above Median Dummy			0.09		0.10			0.01		0.02
(Delinquency Rate: $\geq$ 9 Months)			(0.07)		(0.08)			(0.06)		(0.06)
Interest Rate X Above Median Dummy			-0.03		-0.04			0.05		0.06
(Proportion of Secured Loans)			(0.06)		(0.07)			(0.06)		(0.06)
Interest Rate X Above Median Dummy			0.01		0.01			-0.04		-0.05
(Deposit Rate: Total Accounts)			(0.11)		(0.12)			(0.10)		(0.10)
Interest Rate X Above Median Dummy				-0.13	-0.11				0.17*	0.22**
(Leverage Ratio)				(0.12)	(0.12)				(0.10)	(0.11)
Specification	IV	IV	IV	IV	IV	IV	IV	IV	IV	IV
Lewis and Mertens F-Statistics	74.50	71.04	32.22	55.68	17.63	74.50	71.04	32.22	55.68	17.63
Branch FE	X	X	X	X	X	X	X	X	X	X
Bank X Quarter FE	X	X	X	X	X	X	X	X	X	X
District X Quarter FE	X	X	X	X	X	X	X	X	X	X
Baseline Mean	4.33	4.33	4.33	4.33	4.33	2.58	2.58	2.58	2.58	2.58
Number of Banks	39	39	39	39	39	39	39	39	39	39
Number of Branches	1855	1855	1855	1855	1855	1855	1855	1855	1855	1855
Observations	22260	22260	22260	22260	22260	22260	22260	22260	22260	22260

Notes: This table reports results from estimating the IV Regression (6) via TSLS with additional terms interacting the endogenous variable, the branch-level average interest rate, with precap branch-level characteristics. We extend the analysis of Table 4 by considering the extensive margin (log number of outstanding loans) and intensive margin (log average outstanding amount) as the dependent variables. We instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Likewise, we instrument each interaction term with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$ interacted with the branch-characteristic variable. We define population density and lending HHI as in Table 1. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. We define the leverage ratio as the 2008 year-end ratio of total debt outstanding to total assets at the parent-bank level. The final row refers to the interaction term between interest rates and the nine-month delinquency rate at the parent-bank level. For each specification, we report the Lewis and Mertens (2022) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instruments. *p<0.1; **p<0.05; ***p<0.01.

D Event-Study Effects of Caps on Other Credit Market Outcomes

In this appendix, we present event-study results for the outcomes presented in Section 4.5 for other non-lending credit market outcomes, including delinquency rates, collateralization, deposit amounts and rates, and the sectoral composition of lending. Mirroring the IV results, we uncover no statistically significant reduced-form effects on these outcomes. The only exception is the sectoral shares (Table D.3), which shows statistically significant effects for 2010Q1-Q4. However, the magnitudes are negligible relative to the baseline mean, which is consistent with a precise, yet insignificant IV effects in Table 7 of our paper.

Table D.1: Event-Study Effects of Cap-Induced Changes in Interest Rates on Delinquency and Collateralization

	Delinquency Rate: $\geq$ 9 Months	Delinquency Rate: $\geq$ 3 Months	Proportion of Secured Loans	Predicted Delinquency Rate
	(1)	(2)	(3)	(4)
TrtIntensity x 2008Q1-Q4	0.001	-0.001	0.002	-0.001
	(0.002)	(0.003)	(0.011)	(0.001)
TrtIntensity x 2009Q2-Q4	-0.000	-0.000	-0.007	-0.001
	(0.002)	(0.004)	(0.008)	(0.001)
TrtIntensity x 2010Q1-Q4	-0.001	-0.000	0.003	-0.000
	(0.003)	(0.004)	(0.009)	(0.001)
Specification	OLS	OLS	OLS	OLS
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	0.044	0.056	0.641	0.057
Number of Banks	39	39	39	39
Number of Branches	1855	1855	1855	1855
Observations	22260	22260	22260	22260
Adj. R-squared	0.757	0.735	0.722	0.957

Notes: This table reports results from estimating the reduced form regression (5) for 9-month and 3-month delinquency rates, the fraction of loans secured by physical collateral, and the predicted delinquency rate. The predicted delinquency rate is defined as the weighted average of bank-sector delinquency rates, using the outstanding loan amounts by sector as weights. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

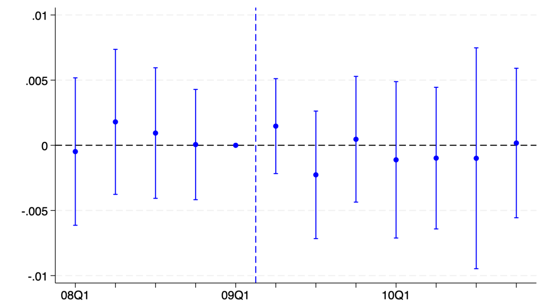
Table D.2: Event-Study Effects of Cap-Induced Changes in Interest Rates on the Deposit Market

	Deposit Rate: Total Accounts	log Deposit Amount: Total Accounts	Deposit Rate: Corporate Accounts	log Deposit Amount: Corporate Accounts
	(1)	(2)	(3)	(4)
TrtIntensity x 2008Q1-Q4	0.01	-0.02**	-0.00	0.07
	(0.01)	(0.01)	(0.01)	(0.07)
TrtIntensity x 2009Q2-Q4	-0.01	0.01	-0.02	-0.01
	(0.01)	(0.01)	(0.02)	(0.06)
TrtIntensity x 2010Q1-Q4	0.00	-0.00	0.00	-0.01
	(0.01)	(0.01)	(0.01)	(0.08)
Specification	OLS	OLS	OLS	OLS
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	5.91	6.49	5.93	4.66
Number of Banks	39	39	39	39
Number of Branches	1855	1855	1712	1712
Observations	22260	22260	13035	13035
Adj. R-squared	0.97	0.95	0.93	0.71

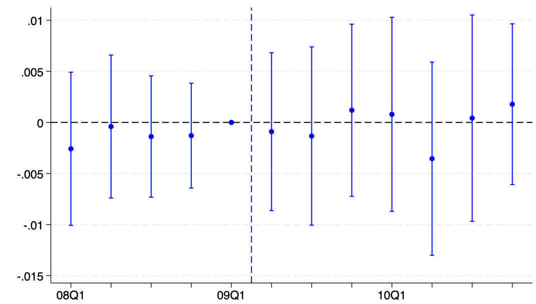
Notes: This table reports results from estimating the reduced form regression (5), where now the set of outcome variables includes deposit rates and deposit amounts across all types of accounts (corporate plus individual) and just corporate accounts. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

Figure D.1: Impact of Interest Rate Cap on Deposit Rates, Loan Performance, and Collateralization

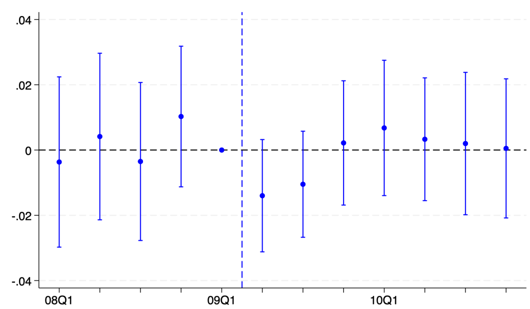
(a) **Delinquency Rate: $\geq$ 9 Months Overdue**



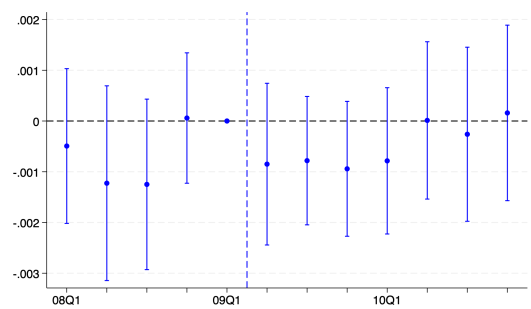
(b) **Delinquency Rate: $\geq$ 3 Months Overdue**



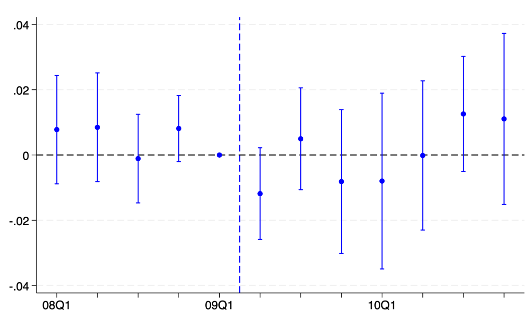
(c) **Proportion of Secured Loans**



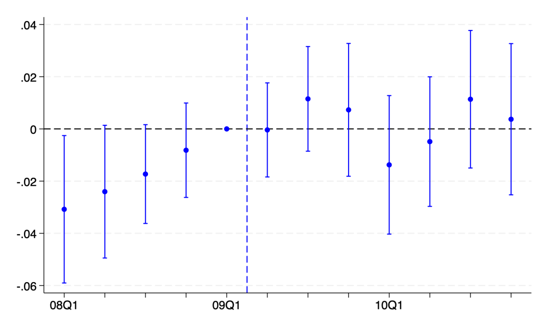
(d) **Predicted Delinquency Rate**



(e) **Deposit Rate**



(f) **log Total Deposit Amount**



Notes: The figure plots the estimated coefficients of treatment intensity interacted with quarter dummies from event-study equation (5) with proxies for lenders' costs of supplying credit as the outcomes, including: delinquency rates, the proportion of loans secured by physical collateral, predicted delinquency rate, annualized deposit rates on individual accounts, and the total deposit amount for individual accounts. We omit the quarter before the cap reform (2009Q1) as the reference category.Bangladesh Bank announced the interest rate cap on April 19, 2009 (2009Q2), with the cap effective immediately. 95% confidence interval bars obtained from clustering standard errors at the branch level.

Table D.3: Event Study Effects of Cap-Induced Changes in Interest Rates on Sectoral Composition

	Share of Outstanding Loans: Manufacturing	Share of Outstanding Loans: Import or Export	Share of Outstanding Loans: Others
	(1)	(2)	(3)
TrtIntensity x 2008Q1-Q4	0.004	0.004	-0.008
	(0.007)	(0.006)	(0.006)
TrtIntensity x 2009Q2-Q4	0.004	0.003	-0.007
	(0.005)	(0.005)	(0.005)
TrtIntensity x 2010Q1-Q4	0.014*	0.006	-0.019***
	(0.007)	(0.006)	(0.007)
Specification	OLS	OLS	OLS
Branch FE	X	X	X
Bank X Quarter FE	X	X	X
District X Quarter FE	X	X	X
Baseline Mean	0.165	0.087	0.748
Number of Banks	39	39	39
Number of Branches	1855	1855	1855
Observations	22260	22260	22260
Adj. R-squared	0.830	0.760	0.863

Notes: This table reports results from estimating the reduced form regression (5), where now each outcome is the share of outstanding loans for one of the three broad corporate sector classifications: manufacturing, import or export firms, and other sectors. As before, we instrument the endogenous variable with $TrtIntensity_i \times \mathbb{1}\{t \geq 2009Q2\}$, which captures how much branch i is exposed to the interest rate cap regulation. Robust standard errors clustered at the branch level are in parentheses. In each specification, we include branch, bank-quarter, and district-quarter fixed effects. For each specification, we report the Montiel Olea and Pflueger (2013) heteroskedasticity- and cluster-robust first-stage F-statistic for the excluded instrument. *p<0.1; **p<0.05; ***p<0.01.

E Appendix for Additional Evidence of Ex Ante Market Power

In this appendix, we conduct additional tests for banks' ex ante market power using new branch entry. We start by extending our analysis in Section 5.2 by re-estimating equation (9), repeated below here for convenience, using alternative estimators besides the de Chaisemartin and D'Haultfœuille (2020) estimator we adopt in the main text:

$$Y_{i,t} = \sum_{t=-m,t\neq-1}^{t=n} \xi_t \cdot \text{CompetitorEntry}_{i,t} + \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \varepsilon_{i,t} \tag{E.1}$$

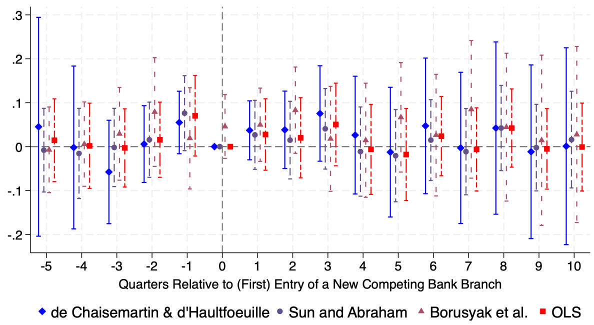
As before, we define $\text{CompetitorEntry}_{i,t}$ as a dummy indicating that the nearest-neighbor competitor bank to the bank belonging to branch i opens a new branch in quarter t in the same subdistrict where i is located. The null effects of entry on branch lending and price setting we document in Figure 7 restrict to the set of branches experiencing competitor entry at some point in the sample period, comparing branches experiencing entry in a quarter to those not yet experiencing entry.

In Figure E.1, we show how the event-study coefficient estimates $\widehat{\xi}_t$ are nearly identical in their magnitude and trend regardless of whether we use OLS or staggered difference-in-differences estimators which either include or exclude never-treated units. This is true across all four of our main branch-level credit market outcomes: interest rates, the log outstanding loan amount, log number of loans, and the log average outstanding loan amount. In this context, a never-treated unit would be a branch never experiencing the entry of a close competitor bank's branch during our sample period. In contrast, the Sun and Abraham (2021) estimator includes only never-treated units in the control group; since in the absence of time-varying covariates the Callaway and Sant'Anna (2021) and Sun and Abraham (2021) estimators will return identical point estimates, we report results for only the Sun and Abraham (2021) estimator. The Borusyak et al. (2024) estimator includes both never-treated and not-yet-treated units in the control group.

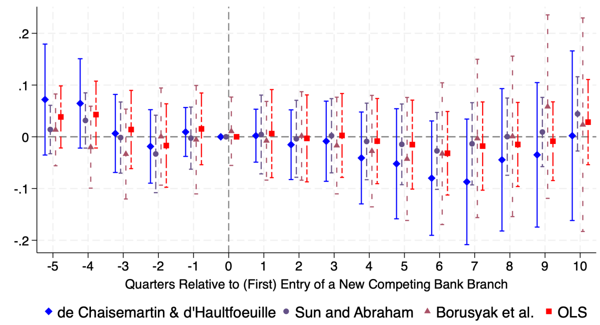
In a second test, as a validation check, we aggregate the data to the subdistrict-quarter level and ask whether the entry of a new branch has a positive effect on total local credit provision while simultaneously having no negative effect on lending by incumbent branches. If true, the interpretation of this hypothesis is that incumbent branches within the same subdistrict face imperfect competition, and therefore the entry of any new branch does not result in them losing

Figure E.1: Null Effects of Entry of Competitor Branch: Alternative Staggered Difference-in-Differences Estimators

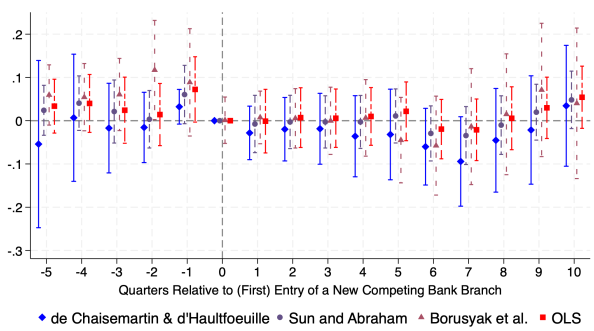
(a) **Interest Rates**



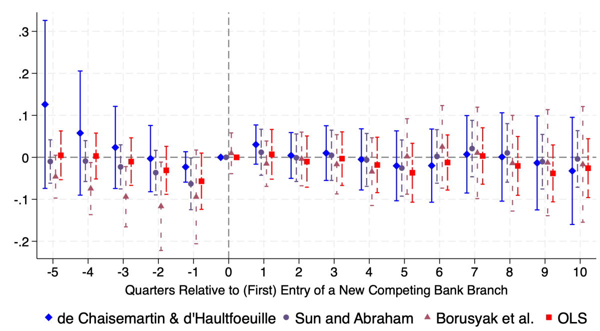
(b) **Log Outstanding Loan Amount**



(c) **Log Number of Loans**



(d) **Log Average Outstanding Loan Amount**



Notes: The figure plots the estimated quarterly event-study coefficients on the sequence of dummies $\text{CompetitorEntry}_{i,t}$ from equation (E.1) indicating the first entry of a branch tied to the nearest-neighbor competitor bank of branch i within the same district. We consider the four main outcome variables adopted throughout the paper: share-weighted average interest rates in Panel (a); log outstanding loan amounts in (b); the log number of loans in (c); and the log average loan size in (d). In each panel, we plot the coefficients from estimating (5) via four estimators: de Chaisemartin and D'Haultfœuille (2020), Sun and Abraham (2021), Borusyak et al. (2024), and OLS. In each regression, we omit the quarter prior to the competing branch's entry as the reference category. 95% confidence interval bars obtained from clustering standard errors at the branch level. We implement the de Chaisemartin and D'Haultfœuille (2020) estimator with 1,000 bootstrap iterations to compute standard errors.

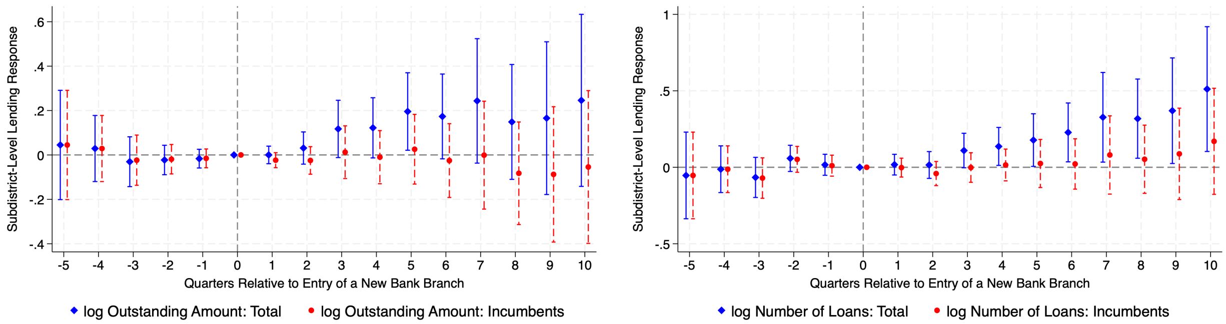
business. We estimate the following regression at the subdistrict level:

$$Y_{Subdistrict,t} = \sum_{t=-m,t\neq-1}^{t=n} \zeta_t \cdot Entry_{Subdistrict,t} + \eta_{Subdistrict} + \psi_{District,t} + \varepsilon_{Subdistrict,t} \quad \text{(E.2)}$$

where now, in contrast to equation (9), the dummies of interest $Entry_{Subdistrict,t}$ are equal to one in

Figure E.2: Effect of New Bank Branch Entry on Total Local Credit Provision

(a) **Log Outstanding Loan Amount** (b) **Log Number of Loans**



Notes: The figure plots the estimated quarterly event-study coefficients on the sequence of dummies $Entry_{Subdistrict,t}$ from equation (E.2) indicating the entry of a new bank branch into a subdistrict. We consider the two main outcome variables adopted throughout the paper: log outstanding loan amounts in Panel (a); the log number of loans in (b). In each panel, we plot the coefficients from estimating (de Chaisemartin and D'Haultfœuille 2020). 95% confidence interval bars obtained from clustering standard errors at the branch level.

the event of *any* new branch entry in the subdistrict-quarter. We then re-estimate (E.2) separately after restricting only to total lending by incumbent branches and compare the estimated $\widehat{\zeta}_t$.

Figure E.2 present the results from this subdistrict-level analysis of the impacts of new branch entry on total lending. Entry of any new branch into a subdistrict has no discernible effect on total lending of incumbent branches, but a positive effect on total lending across all branches in the subdistrict. On the extensive margin, there is a positive effect on credit provision at the subdistrict level, and a statistically null, but positively trending, series of coefficients for incumbent branches. Together, the two sets of tests in this appendix point to banks in Bangladesh facing imperfect competition for new borrowers.

F Testing for Credit Reallocation within the Bank Branch Network

Do banks reallocate credit within their branch network to smooth out the shock to credit demand induced by the interest rate cap by equating marginal loan profitability across branches? Or, do banks respond by reallocating credit provision to less risky segments of the borrower pool or locations where borrowers are easier to screen? The corporate finance literature argues that firms reallocate resources across plants within their internal network in response to shocks to production. Such shocks might include those to local investment opportunities via reduced informational costs (Giroud and Mueller 2015), local consumer demand via house prices (Giroud and Mueller 2019), and place-based corporate income tax incentives (LaPoint and Sakabe 2021). Acharya et al. (2022) demonstrate that banks reallocate residential mortgage loans from the more exposed urban markets to rural areas in response to more stringent leverage limits in Ireland. However, it is an open question whether banks might similarly reallocate credit to reduce their exposure to top-down price regulation such as an interest rate cap.

Unlike firm inputs like capital or labor, since loans are both factors of production and a direct source of revenue for banks, it is not obvious which reallocation channel will dominate when there is an interest rate shock such as a cap and banks respond in a profit-maximizing way. Offering more loans increases revenues from charging interest but also requires banks to retain more capital to cover their position. To isolate credit reallocation across locations within the bank, we estimate the following within-district branch network regression:

$$\begin{aligned} Y_{i,t} &= \sum_{s=-m,s\neq-1}^{s=n} \beta_{1,s} \cdot TrtIntensity_i \times \mathbb{1}\{t=s\} \\ &+ \sum_{s=-m,s\neq-1}^{s=n} \beta_{2,s} \cdot \sum_{k\in District(i)} \omega_k \cdot TrtIntensity_k \times \mathbb{1}\{t=s\} \\ &+ \eta_i + \nu_{Bank(i),t} + \psi_{District(i),t} + \epsilon_{i,t} \end{aligned} \tag{F.1}$$

We augment our main reduced form specification (5) by adding a network effect term capturing the interest rate exposure of i's parent bank to the cap regulation through the internal network of branches other than branch i but within the same district as i. To aggregate individual cross-branch exposures, we take a loan amount share-weighted average of $TrtIntensity_k$, where ω_k are the shares of branch k's lending as a fraction of total bank lending in the district.

For lending as the outcome, if $\beta_{2,s} > 0$, then banks smooth out the interest rate cap shock across branches in their network. However, if $\beta_{2,s} < 0$, then banks instead reallocate loans to branches with relatively lower *ex ante* markups. We present results from estimating (F.1) in Table F.1. We find little evidence of banks responding to the rate cap by reallocating loans across branches according to the distribution of *ex ante* markups across their internal network. If anything, banks smooth out the shock to their loan pricing across branches along the intensive margin of credit provision rather than reallocating credit to more infra-marginal branches. The loadings on $\beta_{2,s}$ are positive and marginally significant at the 10% level when we consider log average outstanding amount as the outcome variable.

Table F.1: Effects of the Interest Rate Cap through Bank Branch Network Exposure

	Interest Rate	log Total Outstanding Amount	log Number of Outstanding Loans	log Average Outstanding Amount
	(1)	(2)	(3)	(4)
TrtIntensity x 2008Q1-Q4	0.08	0.00	-0.01	0.01
	(0.06)	(0.04)	(0.03)	(0.02)
TrtIntensity x 2009Q2-Q4	-0.33***	0.09***	0.03*	0.05**
	(0.08)	(0.03)	(0.02)	(0.03)
TrtIntensity x 2010Q1-Q4	-0.45***	0.18***	0.06**	0.11***
	(0.07)	(0.05)	(0.02)	(0.03)
Network TrtIntensity x 2008Q1-Q4	-0.00	0.05**	0.03	0.02
	(0.03)	(0.02)	(0.03)	(0.02)
Network TrtIntensity x 2009Q2-Q4	-0.09**	-0.00	-0.04	0.03*
	(0.04)	(0.02)	(0.03)	(0.02)
Network TrtIntensity x 2010Q1-Q4	-0.09	0.01	-0.04	0.05*
	(0.05)	(0.03)	(0.03)	(0.03)
Specification	OLS	OLS	OLS	OLS
Branch FE	X	X	X	X
Bank X Quarter FE	X	X	X	X
District X Quarter FE	X	X	X	X
Baseline Mean	14.86	6.96	4.32	2.64
Number of Banks	38	38	38	38
Number of Branches	1596	1596	1596	1596
Observations	19152	19152	19152	19152
Adj. R-squared	0.87	0.96	0.91	0.93

Notes: The table presents results from estimating equation (F.1), which relative to our baseline reduced form specification includes additional interaction terms of *TrtIntensity* with a within-bank-district network exposure term. We define branch i's network exposure as the leave-one-out loan share-weighted average of *TrtIntensity* across all other branches $k \neq i$ but located in the same district as branch i. Robust standard errors clustered at the bank level. *p<0.1; **p<0.05; ***p<0.01.