

# KNOW THYSELF: ACCESS TO OWN CREDIT REPORT AND THE RETAIL MORTGAGE MARKET<sup>\*</sup>

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## Abstract

Borrowers may misestimate their probability of mortgage approval in the absence of precise signals of creditworthiness. Credit reports, which contain such signals, became easily accessible for all U.S. consumers since 2005, while it was already the case in seven states. A difference-in-differences strategy exploiting this change shows that pool quality of mortgage applicants improved as a result—approvals increased and subsequent delinquencies decreased. These findings are consistent with a mechanism where under-estimators enter the applicant pool and over-estimators drop out, because easier access to credit reports reduces misestimation of one's own probability of mortgage approval. Additional findings rule out supply-driven explanations.

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**Keywords:** Credit Reports, Information Provision to Consumers, Household Finance, Mortgages, Regulation of Credit Information

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*More than half of potential homebuyers did not apply for mortgage because they feared they would be rejected, putting their dream of owning a home on pause. Among the ones who feared, nearly three quarters admitted they haven't taken any steps to find out if they qualify for a mortgage.*

*—LoanDepot LLC national survey, 2014*

Borrowers may misestimate their probability of mortgage approval if they do not know about their own creditworthiness. Some may overestimate their chances. This may be why many consumers get rejected due to credit history even though it can be checked before applying. In fact, from 2000 to 2008, twice more applications were rejected due to credit history than were due to debt-to-income (DTI) ratio. Others may underestimate their chances and end up not applying for mortgages despite needing it, because they think they would be rejected (discouraged borrowers). About 13% of U.S. households report as being discouraged in their mortgage decisions ([Survey of Consumer Expectations, 2013–2020](#), SCE).

The economic consequences of such misestimation may be large. The over-estimators incur rejection costs which they could have avoided had they not applied. These costs include an increase in the probability of rejection of all future credit applications and a potential rise in the interest rates. The under-estimators forsake the opportunity of achieving the American dream of owning a home, an opportunity some may have got if they checked their credit reports. Evidence suggests that while lenders almost always use information in credit reports to decide on credit applications, only about 8.4% credit-using consumers in the early 2000's accessed the reports ([Nott & Welborn, 2003](#), p.9).

Motivated by these facts, this paper asks: Is there a link between ease of access to credit reports for consumers and mortgage market outcomes? Specifically, how does easier access to credit reports for consumers affect the mortgage demand, origination, and repayments, and what are the underlying mechanisms?

The primary finding is that easier access to credit reports improves mortgage market outcomes. Specifically, approval ratio increases and delinquencies decline subsequently, suggesting an improvement in the pool of mortgage applicants. A consumer *self-learning* mechanism appears to be at work. Those whose reports signal high approval probabilities may apply for credit (i.e., enter the applicant pool), while those whose reports signal otherwise may either apply to a subprime lender, or may not apply for credit (i.e., drop out of the pool). This consumer-driven sorting results in a better pool of potential borrowers, leading to an increase in approvals. At the same time, the sorting may increase or decrease the demand for credit depending on the

prior distribution of the over- and under-estimators of approval probabilities and those who may be unaware of the role the reports play in a credit assessment process.

The federal *Fair and Accurate Transaction Act* (FACTA) was enacted in 2003 providing all U.S. consumers a right to access three free credit reports for free annually. Subsequently, in 2005, [www.annualcreditreport.com](http://www.annualcreditreport.com) website was established from where consumers could access the free reports in just a few clicks (Figure I). Prior to thus, consumers had to go through many procedural hurdles to access the reports. They needed to call or write a letter to the credit reporting agencies (CRAs) and pay about USD 8–9 to make a request for the reports.<sup>1</sup> If they succeeded in making the request, the reports could still take a week or more to arrive ([Golinger & Mierzwinski, 1998](#)).

Seven states—Colorado, Georgia, Maine, Maryland, Massachusetts, New Jersey, and Vermont—had enacted local laws allowing the state’s residents to access free credit reports before FACTA was enacted.<sup>2</sup> The local laws in these *early* states had driven the usage of credit reports many times higher than the usage in other states. According to one study, the usage rate was 2 to 4 times higher in these states than the rest before FACTA ([United States Senate, 2004](#), p. 506). Other evidence suggest that, relative to the national average, the usage of credit reports was 250% higher in GA, 204% higher in MD, 153% higher in CO, 35% higher in NJ, and 25% higher in MA ([United States Senate, 2004](#), statement of Senator Bennett, p. 376).

Higher usage of credit reports in the early states appear to be a result of the local regulations related to consumer credit. Colorado residents, for example, would receive a notification whenever a negative piece of information was to be added to their credit reports. The notification also reminded them about their right to obtain a free credit report and provided a letter to obtain the reports ([United States Senate, 2004](#), p. 564). Georgia had provisions of two free credit reports annually. In effect, procedural hurdles in obtaining the reports likely explains the differential usage of the reports across the early states and the rest.

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<sup>1</sup> The three national CRAs—Equifax, Experian, and Transunion—blocked calls of millions of consumers who wanted to discuss the content of their credit reports. In 2000, the CRAs settled a lawsuit brought against them for this action ([Federal Trade Commission, January 13, 2000](#)). Moreover, the trivial USD 8 could be paid either through a credit card over the phone or through a bank cheque mailed with the request letter, hence the consumers with no access to credit card or formal banking faced higher procedural hurdles (about 50% of Black and Hispanic households and 20% of non-hispanic white households, [Federal Reserve Board of Governors \(2007, P. 135, Table 4\)](#)).

<sup>2</sup> The legislative bills and the timing of enactment of the local free credit report laws for each of these states are as follows: CO in 1997 through Senate Bill (S.B.) 133; GA in 1996 through House Bill (H.B.) 1632; MD in 1992 through S.B. 20; NJ in 1997 through Assembly Bill (A.B.) 2787, enacted as New Jersey Fair Credit Reporting Act; MA in 1995 through S.B. 79; VT in 1992 through S.B. 453; and ME in 2003 through H.B. 419. A state resident is a person domiciled in the state or who maintains a permanent place of abode; and spends in the aggregate more than six months of the taxable year in the state.

To examine the causal effect of easier access to credit reports for consumers and the mortgage market outcomes, this paper exploits the pre-existing difference in the usage of credit reports in the states which already had free credit report laws and the states bordering them that did not have such laws. A single-event difference-in-differences (DID) design is set up in which the control group consists of all the early states except Maine (since the local law in Maine and FACTA were enacted in the same year). The six control states are referred to as the pre-FACTA states. The treatment group consists of all the 20 states bordering the six control states. In effect, *late-treated* states serve as the treatment group whereas *early-treated* states serve as the control group. The DID estimator is the two-way fixed effects (TWFE) estimator.<sup>3</sup> The event year is 2005, the year in which the website was established, and sample period runs from 2000 to 2008.

In a bid to sweep out the confounding effects of the local economic conditions from the treatment effect, the sample is restricted to only the counties at the border between the treated and control states, similar in spirit to the empirical strategy of [Huang \(2008\)](#) and [Dube, Lester, and Reich \(2010\)](#). The key outcome variables are analyzed at the census tract level, a sub-county micro area that roughly encompasses a population of just about 4,000. Incorporating fixed effects at the census tract level further removes the effect of time-invariant differences across the states, such as recourse or non-recourse mortgages and judicial or non-judicial foreclosures.

The issue of endogeneity in the assignment of the treatment and control groups is considerably alleviated in this empirical strategy. While the “treatment” is effected by FACTA, and it was binding for all 50 states in 2005; the “control” is effected by the local state laws, and it occurred over a period of time well before 2005. Thus the treated states did not opt to become treated, they were mandated to do so. Moreover, the FACTA enactment in 2003 does not appear to be an endogenous response to the prevailing economic conditions. It was enacted to perpetuate the provisions of an existing federal law, the *Fair Credit Reporting Act of 1970* (FCRA), which was set to expire in 2003 (via its amendment in 1996) ([Nott & Welborn, 2003](#)). Hence, FACTA was essentially a repackaged FCRA with a key novelty being the annual free credit reports provision. Since other FACTA provisions already existed under FCRA, they do not contaminate the treatment effect.

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<sup>3</sup> Since this design is not based on a staggered treatment, but on a one-shot treatment, and since the early-treated states were treated deep in the past, outside the sample period, the framework of [Goodman-Bacon \(2021\)](#) suggests that the negative-weights issue of TWFE highlighted in [Borusyak, Jaravel, and Spiess \(2021\)](#), [De Chaisemartin and d’Haultfoeuille \(2020\)](#), and [Sun and Abraham \(2020\)](#) may not arise. Further discussion on this issue appears in Section (2), Empirical Research Design.

Consumers displayed considerable interest in acquiring free credit reports through the website. Anecdotal evidence suggest the website issued about 52 million credit reports in the first two years ([Wikipedia, n.d.](#)). Given that about 190 million individuals in the U.S. had active credit in 2003 ([U.S. House of Representatives, 2003, p.6](#)), 14% of them accessed credit reports yearly from the website alone, a 66% increase over the 8.4% usage in the pre-event period. The act also appears to have raised general awareness about the reports. Interest in free credit reports, measured using Google Search Interest for the keyphrase “Free Credit Reports”, heightened at the time website was established.

The primary finding is that making the reports easier for consumers to access resulted in an increase of 1 percentage point in the approval ratio and 13.8%–16.0% in the number of mortgage applications, and the effect is concentrated in owner-occupied mortgages. The equivalent dollar amount aggregated across the treated bordering counties is about USD 5.5 billion due to the former effect and about USD 38.1 billion due to the latter. The increase in approval ratio is consistent with the self-learning mechanism that predicts that the applicant pool should improve. Furthermore, the higher number of applications indicates that mortgage borrowers on average tend to underestimate their probability of approval, a contrasting finding to the common belief that consumers tend to be overconfident.

Next, the trends in mortgage delinquencies across the two areas are analyzed. If an improved applicant pool underlies the increase in origination, the delinquency rate should decrease, or at least not increase. Indeed, relative to the mortgages from the control areas, those from the treated areas originated in the event-year were *less* likely to become delinquent, but those originated in the pre-event year were just as likely to become delinquent.

Considerable evidence points to the self-learning mechanism. First, the paper assumes that the areas where pre-event rejection ratio due to credit history was high but due to DTI was low are likely the areas where borrowers tend to overestimate the approval probabilities. The mechanism predicts that these areas should experience fewer increase in applications than the rest (since over-estimators drop out). The increase in application was indeed significantly less in the census tracts that were in the highest tertile of the pre-event rejections due to credit history but in the lowest tertile of rejections due to DTI. Second, mortgage-related cognizance among the borrowers seems to increase. Specifically, the treated areas saw a statistically significant decrease in credit-history related mortgage rejections in the *ex-ante* high rejection areas but no decrease in DTI-related denials. This points to an increased learning among borrowers about

their credit history. Also, the fraction of total applications withdrawn while being processed dropped, indicating a reduced tendency to formally apply to multiple lenders and hence saving the costs of multiple applications.

It is also important to understand which borrowers benefit more from easier access to credit reports. The areas with an *ex-ante* high creditworthiness saw a greater increase in approval ratios and mortgage applications. This is in line with the idea that the reports aid consumers in assessing their creditworthiness. Then, for the lowest-income-quartile borrowers, the approval ratios increased statistically significantly in the treated areas vis-à-vis the control areas, but the number of applications did not. Since lower income is associated with creditworthiness overestimation (Perry, 2008), the self-learning mechanism predicts that the correction for the overestimation for these consumers will result in fewer applications (drop out), leading to what the regression estimates suggest—a rise in approval ratios but not in number of applications. Third, treated areas saw proportionally more first-time homebuyers.

Finally, a host of tests examining the response of lenders rule-out the supply-side explanations for the increase in origination. Specifically, relative to the control areas, the treated areas saw (i) increased mortgage interest rates; (ii) the high-lenders-density areas did not see more origination or approvals vis-à-vis the low-lender-density areas; and (iii) private securitization of mortgages did not significantly increase. Furthermore, banks that had larger mortgage origination in the treated areas in the pre-event year saw higher financial performance post event, and those with higher pre-event liquidity originated more mortgages post-event.

All in all, the conclusion is that making it easier for consumers to access their credit reports brings about changes in the mortgage market that are indicative of an improvement in the applicant pool. Given that the findings are causal, policies aimed at encouraging consumers to check their credit reports and educating them about its role in credit approval may yield similar improvements, not just for the mortgage markets, but for any consumer credit market.

This paper primarily relates to the literature on effects of information provision on credit market participants. This is the first paper to show that making it easier for consumers to access credit reports leads to improved mortgage market outcomes in a manner consistent with an improved applicant pool. A field experiment by Homonoff, O'Brien, and Sussman (2019) reveals that borrowers are less likely to default when provided with information on their FICO® scores. Similarly, Mikhed (2015) shows that borrower participation in a free FICO scores program is associated with lower delinquencies, reduced credit utilization, and increased credit

card spending. Using a pair of policy changes in Chile, [Kulkarni, Truffa, and Iberti \(2018\)](#) find that increasing disclosure about financial products leads to lower defaults for sophisticated borrowers, and that standardizing financial products leads to lower defaults for unsophisticated borrowers. Also, bankruptcy flag removal from the credit reports raises mortgage borrowing by consumers ([Dobbie, Goldsmith-Pinkham, Mahoney, & Song, 2016](#)), and lowers the cost of credit for poorer defaulters and increases it for poorer non-defaulters ([Lieberman, Neilson, Opazo, & Zimmerman, 2018](#)).

This paper also speaks to the extensive literature on household financial literacy. Low financial literacy leads to mortgage delinquencies and foreclosures ([Gerardi, Goette, & Meier, 2010](#)), poor mortgage choice ([Moore, 2003](#)), large debt ([Lusardi & Tufano, 2009](#); [Stango & Zinman, 2009](#)), and lower ability to benefit from loan-modification contracts when in distress ([Hundtofte, 2017](#)). On the other hand, educational intervention improves consumers' financial product purchases ([Balakina, Balasubramaniam, Dimri, & Sane, 2020](#)). This paper shows that making it easier for consumers to access their credit reports lowers mortgage delinquencies and raises mortgage application approval ratios and mortgage-related cognizance among borrowers.

The rest of this paper is organized as follows. Section (1) describes the U.S. laws related to consumers' access to credit reports, Section (2) presents the research design, and Section (3) describes the data this paper uses. Section (4) discusses the main results, and Section (5) contains supplementary results that aid interpretation of the main findings. Finally, Section (6) concludes the paper.

## 1 U.S. Laws Governing Consumers' Access to Credit Reports

Enacted in 1973, the FCRA was the first legislation regulating the information credit reporting agencies collect and the manner in which consumers could access it. The act provided consumers the right to see the contents of their credit reports, except for the credit score, under specific yet restrictive circumstances. Consumers could receive a free report if they made a request within 60 days after receiving a notice of an *adverse action* taken against them on the basis of the information in the report ([Nott & Welborn, 2003](#)).<sup>4</sup> An amendment to the FCRA in 1992

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<sup>4</sup> An adverse action notice can be sent to a consumer by the *user* of a consumer report (e.g. banks, financial institutions, insurance firms) or a debt collection agency affiliated with the CRA stating that the consumer's credit rating may be or has been adversely affected. Under the FCRA, a consumer can also receive credit report free of charge once in 12 months by making a request to the CRA certifying that she/he either: (A) is unemployed and intends



further mandated that the cost of disclosure of credit information should be reasonable, and the next amendment in 1996 capped the cost of the disclosure at USD 8 and provisioned that the FCRA would expire in 2003.<sup>5</sup>

It was when the FCRA was to expire that FACTA was enacted, specifically to perpetuate the FCRA's existing provisions while also adding the new annual free credit report provision. FACTA was signed into law on December 4, 2003, and *inter alia* it allowed for free annual disclosure of credit reports to consumers by each of the three national credit reporting agencies through a centralized source. Subsequently, the website—[www.annualcreditreport.com](http://www.annualcreditreport.com)—was established in 2005 to distribute the free credit reports.<sup>6</sup> Panel (A) and (B) of Figure(I) shows the homepage and the “frequently asked questions” section of the website, respectively and Figure(II) shows the summary page of a credit report obtained from the website.

Notwithstanding the federal regulations on consumer credit reporting, seven states (CO, GA, MA, MD, NJ, and VT) enacted local state laws over 1992 to 2003 that allowed their residents to access free credit reports (see Footnote (2) for details of the enactments). For example, Colorado enacted its free credit report law on April 21, 1997 through Senate Bill 133. Section 4, paragraph (E) of this bill added the following to Title 12 Article 14.3-104 of the Colorado Statute:

*(E): Each consumer reporting agency shall, upon request of a consumer, provide the consumer with one disclosure copy of his or her file per year at no charge whether or not the consumer has made the request in response to the notification required in paragraph (a) of this subsection.*

These state laws were likely an endogenous response to the local environment. For example, Vermont was the first state to enact its law, in 1992, because TRW (a credit reporting firm, now Experian) in 1991 *mistakenly* recorded the tax bills of each resident of the town of Norwich and 650 of the residents of the town of Woodstock as property tax liens, due to which these con-

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to apply for employment in the 60 day period beginning on the date on which the certification is made; (B) is a recipient of public welfare assistance; or (C) has reason to believe that the file at the agency contains inaccurate information due to fraud.

<sup>5</sup> Even though the cost of credit reports was capped under the FCRA and even though consumers could access free credit reports under specific circumstances, only a small fraction of credit-using consumers did so. Out of approximately 1 billion credit reports generated annually, only 1.6% were disclosed to consumers (Avery, Calem, & Canner, 2004). Only 5.25% of these were proactively requested by consumers, while 94.75% were disclosed following the various FCRA provisions (Nott & Welborn, 2003). Thus, only 0.084% of all credit reports were generated as a result of a consumer request.

<sup>6</sup> The website was rolled-out in four phases over nine months, from December 2004 to September 2005. Phase (I) rollout was on Dec 1, 2004 in 13 states: AK, AZ, CA, CO, HI, ID, MT, NV, NM, OR, UT, WA, and WY. Phase (II) rollout was on March 1, 2005 in 12 states: IL, IN, IA, KS, MI, MN, MO, NE, ND, OH, SD, and WI. Phase (III) rollout was on June 1, 2005 in 11 states: AL, AR, FL, GA, KY, LA, MS, OK, SC, TN, and TX. Phase (IV) rollout was on September 1, 2005 in the remaining 14 states and Washington D.C. CT, DE, DC, ME, MD, MA, NH, NJ, NY, NC, PA, RI, VT, VA, and WV.



sumers would have been rejected for credit by every lender had they requested it ([Associated Press News, December 24, 1992](#)). In the same year, Maryland enacted a similar law.

## 2 Empirical Research Design

As discussed, this paper uses a DID setting in which six pre-FACTA states—CO, GA, MA, MD, NJ, and VT—constitute the control group and the states bordering these constitute the treatment group. Panel (A) of Figure (III) shows these states on the map of the contiguous U.S. The regression sample focuses only on those counties that lie at the borders between the treated and control states, and Panel (B) of Figure (III) shows these counties on the map of the contiguous U.S. The event year is 2005, the establishment year of the website. The sample period is chosen to be from 2000 to 2008 to allow for sufficient post-event observations.

This empirical design is influenced, but not compromised, by the fact that the enactments of local laws are endogenous. This is because this design does not utilize the staggered adoption by the six control states, but rather relies on the federal enactment of FACTA after the staggered adoption by those states and the fact that FACTA was binding on all the states.

Differences in state regulations on housing and mortgages—e.g., recourse versus non-recourse mortgages ([Ghent & Kudlyak, 2011](#)) and judicial versus non-judicial foreclosures ([Gerardi, Lambie-Hanson, & Willen, 2013](#))—may introduce effects that confound the treatment effects. Additionally, credit-related regulations that are enacted within the sample period, such as the adoption of Anti-predatory Lending laws (APL) by twenty U.S. states over 2000 to 2006 ([Di Maggio & Kermani, 2017](#)), may further aggravate the issue.<sup>7</sup>

However, the DID design makes the estimates robust to any state-level differences that do not change over the sample period, such as the recourse versus non-recourse mortgages and judicial and non-judicial foreclosures. The confounding effects of staggered regulatory adoption, too, get averaged out in the estimation when the timing of adoption and the states who adopt them are different from the timing of the natural experiment and the treated and control states. This issue is further alleviated by the fact that removing each of the control and associated treated states one at a time yields similar estimates for key outcome variables (discussed in the Results section).

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<sup>7</sup> Table (A1) in the Online Appendix lists the treated and control states and their status with respect to these regulations. The distribution of these regulatory differences is as follows: 90% of the treated and all the control states have recourse mortgages; 45% of the treated and 83.3% of the control states have judicial foreclosures; and 35% of the treated states and 83% of control states adopted the APL laws within the sample period.

The contiguous-county design across state borders also alleviates the potential confounding effects of local idiosyncratic trends, since such trends are not likely to vary widely across neighboring areas and macroeconomic shocks affect neighboring areas roughly at the same time (Dube, Lester, & Reich, 2016), making this design one of the most compelling identification strategies (Allegretto, Dube, Reich, & Zipperer, 2017).<sup>8</sup> A similar empirical approach has been used in Huang (2008) and Dube et al. (2010).

## 2A The DID estimator

The estimator used in the paper is the two-way fixed-effects (TWFE) estimator, specified as

$$Y_{icsjt} = \beta_0 + \beta_1 \times \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}, \quad (1)$$

where  $Y_{icsjt}$  is the outcome variable measured in year  $t$  for a census tract  $i$  from a county  $c$  lying at the border between treatment state  $s$  and control state  $j$ .  $t$  ranges from 2000 to 2008, and  $j$  ranges from one to six, corresponding to each of the six control states.  $\text{Post}_t$  takes value 0 for year  $t < 2005$  and value 1 for year  $t \geq 2005$ .  $\text{Treat}_{icsj}$  is 0 for all the census tracts  $i$  in counties  $c$  from control states  $j$ , and it is 1 for all the census tracts in counties from the treatment states  $s$ .  $\beta_1$ , the coefficient of interest, captures the treatment effect, which is the change in the dependent variable in the treated counties relative to the control counties occurring in the post-event period relative to the pre-event period. Standard errors are clustered at the county level to provide for correlation in error terms for the observations from census tracts belonging to the same county.

*Economic controls* in the equation represent the co-variates. They include a host of time-varying county- and state-level variables capturing local economic and credit conditions. These are the number of mortgage lenders (in log) in a census tract and the annual growth rate of county's income per capita, county's aggregate employment and state's gross domestic product (GDP). To ensure that the treatment effects are not influenced by the co-variates, all regressions are estimated both with and without the co-variates.

$\alpha_i$  represents *Census Tract* fixed effects, the first of the two-way fixed effects. These account for any time-invariant differences across census tracts at a highly granular geographic area that

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<sup>8</sup> In the context of this paper, another viable approach is to use a synthetic control matching procedure. However, this method, too, may place greater weights on *nearby areas* (Allegretto et al., 2017) as it replicates the unobserved counterfactual by taking a weighted average of the observable units. A benefit of a contiguous-county design is that the treatment and control areas map one-to-one to observable geographic areas, enabling tighter links to the real-world data, while a disadvantage is that economic environments across states are inherently different.

encompasses a population of just about 4,000. As census tracts are smaller geographic areas than a county or state, these fixed effects flexibly and fully account for any state-level time-invariant differences, including recourse and non-recourse mortgages, judicial and non-judicial foreclosures etc.

$\gamma_{j,t}$  represents *Border*  $\times$  *Year* fixed effects, the second of the two-way fixed effects. Here  $j$  refers to the border of a control state  $j$ . These fixed effects are formed by the interaction of the border of each of the control states with year, allowing for each region to have its separate time trend (where a region consists of the counties at the border between a given control state and all contiguous states). Thus any regional shock that may affect the regions across different years are flexibly and robustly accounted for.<sup>9</sup>

All in all, the two fixed effects and the time-varying economic controls are expected to reasonably account for confounding effects of local economic conditions on the outcomes of interest, and thus should allow for cleaner estimation of the treatment effects.

## 2B Is the TWFE an appropriate estimator for the current DID design?

A key issue with the TWFE estimator is that in *staggered* DID designs, it may aggregate individual treatment effects by assigning “negative weights” to some of them (Borusyak et al., 2021; De Chaisemartin & d’Haultfoeuille, 2020; Sun & Abraham, 2020). Since the estimator is the variance-weighted average of the treatment effects, the negative weights occur in staggered designs when the treatment effects are heterogeneous across time and/or the treated units (Goodman-Bacon, 2021). Here the issue of heterogeneous treatment effects across time does not arise, since the current paper uses a *single-treatment* DID design, not a staggered treatment.<sup>10</sup> The other issue of the treatment effect being heterogeneous across treated units remains a noteworthy limitation. However, the results in the paper remain largely supportive of the conclusions, since the key estimates are robust across the sub-samples formed by removing, one at a time, each of the control states and the respective contiguous treated states.

Time-varying co-variates also potentially introduce bias in the estimator (Goodman-Bacon, 2021), but the conclusions of this paper are robust to this issue, as all the estimates are quali-

<sup>9</sup> Consider the control state CO. All census tracts from the counties at the border between CO and the surrounding states—WY, UT, AZ, NM, OK, KS, and NE—take the same value for  $j$ , and thus are grouped as one unit (region). Thus the fixed effects only utilize the variation *within* each of the six such regions.

<sup>10</sup> The staggered adoption of the local laws by the control states over 1992–1997 is not an issue either, since the post-event outcomes are measured at least eight years after the last adoption by a control state. With smaller gaps, the longer-lasting dynamic treatment effects *may* affect the outcome variable in the control states and the later-treated states at the same time, compromising the parallel-trends assumption.

tatively and quantitatively similar, *either with or without* the co-variates. Finally, the TWFE also requires random assignment of the treatment. Though the assignment is not carried out randomly, the timing and circumstances of the FACTA enactment are unrelated to the states' actions, and thus the requirement is expected to be largely satisfied.

In the end, the TWFE relies on the parallel-trends assumption: the treated states would have had trends similar to that of the control states in the absence of the treatment. Though the assumption is unverifiable, Panel (A) of Figure (IV) plots the trend of the mean approval ratio across the two groups before the event, and they seem to be parallel. Furthermore, Panel (B) of Figure (IV) plots the coefficients ( $\beta_k$ ) from regression of the *Approval Ratio* according to the following specification:

$$Y_{icsjt} = \beta_0 + \sum_{k=T-3}^{T-1} \beta_k \text{Treat}_{icsj} \times \text{Event}_k + \sum_{k=T+1}^{T+4} \beta_k \text{Treat}_{icsj} \times \text{Event}_k + \alpha_i + \gamma_{j,t} + \varepsilon_{icsjt}, \quad (2)$$

where  $\text{Event}_k = 1$  if  $t = T - k$ ,  $\text{Event}_k = 0$  if  $t \neq T - k$ ,  $k = \{-3, 4\}$ , and  $T = \text{Event year 2005}$ . The coefficients  $\beta_k$ 's represent the difference in approval ratio for the two groups over the years relative to the pre-event year 2004. We see from the plot in Panel (B) that for the most part no significant difference exists between the treated and control census tracts before the event, but the difference becomes significant afterwards. Overall, the two plots in Figure (IV) reasonably assure that the parallel-trends assumption largely holds.

## 2C Salience of the natural experiment and other validating assumptions

As the central cause that shapes the outcomes of interest are the free credit reports, it should be validated whether consumers had interest in accessing their credit reports, or alternatively, whether the natural experiment was salient for them. The uptick in interest in free credit reports, measured using Search Interest data from Google Trends, strongly suggests so.<sup>11</sup> *First*, the search interest for the key phrase *Free Credit Report* heightened in Jan 2005, coinciding perfectly with the establishment of the website (Panel (A) of Figure (VI)). *Second*, plotting the search interest separately for the treated and control states using the Interest-by-subregion data in Panel (B) of Figure (VI) reveals that, although the interest in free credit reports was simi-

<sup>11</sup> The Google Trends data represent the degree of "search interest" for a given keyword at any time relative to the highest point during the period of analysis over a given region (U.S.). In the time series, a value of 100 represents the peak popularity for the term. A value of 50 means that the term is half as popular. In the cross-section, a value of 100 represents the location with the highest popularity of the keyword as a fraction of total searches in that location. A value of 50 indicates a location that is half as popular. A score of 0 means there were not enough data for this term. Google Trends data start from January 2004.

lar in both sets of states in the pre-event year 2004, the interest was more intense in the treated states relative to the control states after the event.<sup>12</sup> Also, as discussed before, almost 52 million credit reports were issued to consumers through the website in the first two years, representing an increase of 66% in the overall usage of the reports.

One may wonder why making credit reports free affects the market outcomes given that the monetary cost of credit reports has been trivially small at just about USD 8 historically. It appears that the outcomes respond primarily not to the reduction in the monetary cost of the reports, but to the reduction in the procedural hurdles of accessing them. Despite costing so little prior to FACTA, there existed stark differences in the usage of the reports in the states where they were free (control states) and where they were not. Moreover, a mere 0.084% of the approximately 1 billion reports generated annually in the U.S. (in the early 2000's) were proactively requested by consumers (Avery et al., 2004), and the fraction of credit-using consumers who requested their credit reports was just 8.4% (Nott & Welborn, 2003). Furthermore, if one were to assume that all these credit-report-using consumers had applied for a mortgage in 2004, their applications would make up less than 5% of the total number of mortgage applications.

It is also worth pointing out that the credit reports issued under FACTA do not contain the numerical credit score. However, the consumers visiting the website were actively asked if they wished to retrieve their scores from any of the three CRAs. The website not only guided them on how to get the score from within the website itself, but also provided explanations on the role of the reports and scores (see Figure (I)). Hence, it seems reasonable that consumers could understand their creditworthiness better from the website.

Finally, the quantities estimated in this paper measure the true effect conservatively, as the underlying treatment is of the intention-to-treat (ITT) nature—the consumers who wanted to get their credit reports could do that before FACTA as well. These estimates may also be labeled as the *average treated effects on the late-treated* (ATT-LT), because their computation involves comparing the outcomes in the late- and early-treated areas.

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<sup>12</sup> An issue with analyzing cross-sectional trends in the search-interest data (interest-by-sub-region data) is that the data values are normalized by Google within the time interval for which the data are extracted. However, this can be overcome by first extracting the data *separately* for each time interval of interest (one-year intervals, in the current case), and then calculating the mean *separately* within each time-interval for each of the two sets of states. From this plot, it may appear that the popularity in the control states after the event decreased. However, this occurs because the popularity measure is essentially a yearly percentile ranking of states, with 100 being the most popular; so an increase in the rank of one state mechanically decreases the rank of others.

### 3 Data and Summary Statistics

This paper primarily draws on three publicly available datasets: *Home Mortgage Disclosure Act of 1975* dataset (HMDA data) for information on mortgage applications; Federal National Mortgage Agency (Fannie Mae) and the Federal National Home Loan Mortgage Corporation (Freddie Mac) dataset (GSE data) for information on mortgage delinquency performance; and the Call Reports (FFIEC Forms 031/041) for information on financial performance of banks. HMDA dataset provides application-level details on applicants' race and gender, income, loan amount, the financial institution handling the mortgage application, outcome of the application, and geographic location of the property at the census tract level.<sup>13</sup> The GSE dataset contains only mortgages purchased by Fannie Mae and Freddie Mac, and it covers only the 30-year fixed-rate single-family mortgages, the most popular mortgage type in the U.S. The mortgage-level information in these data include the interest rate, DTI ratio, credit score, first-time homebuyer status, investment purpose, and the first three digits of the zip code (zip3) of the mortgaged property. The mortgage performance information include repayment amount and delinquency status, both updated monthly. Finally, Call Reports contain detailed financial information of the U.S. banks.

The detailed steps to process each of these datasets and to link them with one another are provided in Appendix (6), Data Appendix. Mortgages for all purposes and types in the HMDA dataset are included in the sample. The three purposes are—home purchase, refinance, and home improvement, and the three types are—conventional loans, loans guaranteed by Veterans Administration (VA) and Farm Service Agency (FSA)/Rural Housing Administration (RHS), and loans insured by the Federal Housing Administration (FHA). These application-level data are aggregated to the *Census Tract*  $\times$  *Year* panel, resulting in 11,942 census tracts that belong to the bordering counties. There are 7,011 treated census tracts, 4,931 control census tracts, and 89,535 *Census Tract*  $\times$  *Year* observations. Similarly, the mortgage-level GSE data are aggregated to the zip3-state level, leading to 221 unique zip3-states (91 as control and 130 as treated) and 7,599 *Zip3*–*State*  $\times$  *Quarter* observations.

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<sup>13</sup>Until 2003, the census tracts in the HMDA dataset are from the Census 1990 definition, while those from 2004 onward are from the Census 2000 definition. To facilitate the comparison of the tract-level data pre-2003 with post-2003, the census tract-level variables from 2000 to 2003 were scaled using the ratio of population residing in the 1990 tract definition to that in the 2000 definition using data from the [Census Bureau \(2006\)](#). Even though this process is an approximation and introduces some noise in the measurements, it is necessary. The approximation is limited to just 22% of the tracts across the U.S., since 63% of the 1990 census tracts did not see any change across the two censuses and 15% of the 1990 census tracts were wholly combined into various 2000 tracts.

Some other data are also collected from public sources. Survey data on consumers' credit usage are taken from the [Survey of Consumer Expectations \(2013–2020\)](#) Credit Access Survey, a Federal Reserve Bank of New York rotating panel survey fielded since 2013 over the internet every four months. The data on county-level subprime population come from [FRBNY and Equifax \(n.d.\)](#), on state-level economic conditions from the Bureau of Economic Analysis, on census-tract level population characteristics from Census 2000 ([Manson, Schroeder, Van Riper, & Ruggles, 2019](#)), and on county-level employment from the annual survey of County Business Patterns (CBP) ([Census Bureau, 2000–2008](#)).

The key outcome variables are scaled applications and approval ratio. The scaled applications is the number of mortgage applications per 1000 adults in a census tract, and the approval ratio is the ratio of the number of successful applications (action type “1” or “2” in the HMDA dataset) to the number of total applications in a census tract. Similarly, other variables of interest are defined at the census-tract level: the fraction of total applications withdrawn while still under processing and the fraction of total applications denied for credit history or DTI ratio.

Panel (A) of Table (I) shows the summary statistics for the key variables over the sample period. We see that the treated census tracts have fewer scaled applications, lower mortgage approval ratio, and higher denials related to credit history and DTI ratio.<sup>14</sup>

Panel (B) of Table (I) shows the comparison of the treatment and control groups in the pre-treatment period using t-tests for difference in mean, the p-values for which are also shown. Results from the t-tests suggest that the control and treated census tracts differ in pre-treatment years in terms of mortgage-related variables, but are *similar* in the state- and county-level economic characteristics. The similarities in economic characteristics of treated and control counties support the comparison of outcomes across the two groups, whereas the differences in mortgage-related outcomes raise the concern that these groups may also differ on some unobserved characteristics, potentially causing an endogeneity issue. However, since a DID design can accommodate pre-existing differences between the treatment and control subjects so long as they satisfy the *parallel-trends* assumption, the concern is alleviated.

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<sup>14</sup>We see that the four ratios—the approval ratio, two denial ratios, and withdrawal ratio—do not sum to one. There are three reasons for this. First, the reporting of the reason for denial is not mandatory under HMDA regulations; hence an application may be recorded as denied without any stated reason (70.81% of denied applications have at least one stated denial reason). Second, denials could be for reasons other than credit history or DTI ratio. Third, an application might be denied for multiple reasons.



## 4 Results

This section first analyzes credit report usage and the discouraged borrowers phenomenon using recent survey data, followed by the causal effects of easier access to credit reports on the mortgage market outcomes. Results about the self-learning mechanism and heterogeneous effects are presented follows next.

### 4A Baseline Results

#### (I) Survey evidence on credit report usage and discouraged borrowers

A representative survey of U.S. consumers, the SCE Credit Access Survey captures the usage of credit reports and scores among consumers and also their planned credit usage. The rotating panel structure of the survey allows for regression analyses that can accommodate fixed effects and clustering of the standard errors at the  $Year \times Month$  level. Furthermore, sampling weights allow one to make inferences that can be generalized to the population. The data used for this analysis span from 2013 to 2020.

Columns (1) through (3) of Table (II) throw light on the average usage of credit scores and credit reports. To do this, an indicator (dummy) variable for a given characteristic is regressed on a constant. In Column (1) the indicator variable is 1 when a respondent has never checked or requested a credit report and 0 otherwise; in Column (2) it is 1 when the respondent has either never checked his or her credit report or checked it at least more than two years ago (infrequent checkers), and it is 0 otherwise; and in Column (3) it is 1 in when respondent does not know his or her credit score and 0 otherwise. The estimated value of the constant in each case is rather startling: an estimated 8% of the population has *never* checked or requested a credit report (Column (1)), about 20% of the population are infrequent checkers of credit reports (Column (2)), and almost 12% of the population does not know own credit score (Column (3)).

Columns (4) through (6) of Table (II) analyze the phenomenon of discouraged borrowers. One of the questions in the survey asks how likely the respondent is to take out a mortgage and related credit in the next 12 months. Those who are very or somewhat unlikely to do so, or those who assign a less than 10% probability to it are asked for the reason. Defining the indicator variable *discouraged borrowers* as 1 for those who respond “I don’t think I would get approved”, and then regressing it on a constant provides an estimate of the discouraged borrowers. The regression result in Column (4) suggests that among those not planning to

take out mortgages, the fraction of those who are doing so because they are discouraged is about 13%. Furthermore, an indicator for discouraged borrowers is regressed separately on the dummy variable for *infrequent checkers* and for *unawareness* about credit score. Columns (5) and (6) show that the coefficient on both the independent variables is positive and significant, i.e. infrequent checkers and those unaware of their credit score are respectively 3% and 5% *more likely* to be discouraged.

Taken together, these findings imply that not using credit reports and scores is non-trivially prevalent among retail consumers, and this tendency also contributes to potential borrowers becoming discouraged from applying for credit.

## (II) Mortgage market outcomes

The effects of easier access to credit reports is examined on three outcomes: mortgage approval ratio, number of applications per 1000 adults (scaled applications) and house prices. These variables are measured at the census tract level; the regression specification is from Equation (1); and all specifications include *Census Tract* and *Border*  $\times$  *Year* fixed effects. The coefficient of interest is *Treat*  $\times$  *Post*, which estimates the change in the outcome variable in the treated areas relative to the control after the event.

Columns (1) and (2) of Table (III) show the regression results for approval ratio. The specification in the first column is without any co-variables, whereas that in the second includes controls for local economic conditions, namely, the number of HMDA lenders (in log) in a census tract, and annual growth rates of county income per capita, county aggregate employment, and state GDP. Coefficients on *Treat*  $\times$  *Post* suggest that the ratio increased by about 1 percentage point in the treated tracts relative to the control. In dollar terms, keeping the number of applications in the treated areas at the pre-event level, a 1 percentage point increase in approval ratio corresponds to successful mortgages worth about USD 2.75 billion, aggregated across the treated bordering counties.<sup>15</sup> The effect at first may seem trivial, especially since approval ratios are commonly believed to be high, at upwards of 80%, but it is just about 52% in the treated tracts in the pre-event period.

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<sup>15</sup> A 1 percentage point increase in approval ratio is equivalent to  $\sim 2.6$  more successful applications per treated tract ( $96.27$  applications per 1000 adults in the pre-treatment period  $\times 0.01 \times 2.7$  thousand adults per treated tract), about 18,229 more successful applications across the treated bordering counties ( $2.6$  applications  $\times 7,011$  treated tracts), or a  $\sim$ USD 2.75 billion increase in mortgage origination across *all treated tracts from bordering counties* ( $18,229 \times$  USD 150,597 average mortgage amount per application).

Approval ratio increased potentially because borrowers were better informed of their creditworthiness, as the establishment of the website did not affect other aspects of the mortgage process. Specifically, the website did not alter the content of the reports or lenders' access to the reports. It also could not affect any borrower characteristics such as their income, employment, or collateralizability of their assets. However, self-learning mechanism can explain increase in approvals. The website not only made it easy to access credit reports but also provided information on the role of credit reports and scores in credit applications (see the snapshot of the website in Figure (I)). Borrowers possibly learned about their creditworthiness from the reports and sorted themselves better in the market. Creditworthy borrowers stayed or entered the applicant pool whereas those with bad creditworthiness either dropped out of the pool or accessed credit from the subprime lenders, improving the quality of the pool. Later sections examine the mechanism in detail.

The regression results for scaled applications appear in Columns (3) and (4) of Table (III). The applications rose in the treated tracts relative to the control by 13.4–16.6, a 13.9–17.2% increase over the pre-treatment average of 96.3. In real terms, keeping the approval ratio in the treated areas at the pre-event level, the increase roughly translates to USD 37.8 billion, aggregated across the treated bordering counties.<sup>16</sup> The increase in applications indicates that on average consumers tend to underestimate their creditworthiness when it comes to mortgage borrowing.

To ensure that these results are not driven by some specific states, the above regressions are re-estimated over sub-samples formed by removing each of the control states and their respective contiguous treated states one at a time. The coefficients from regressions similar to Equation (1) with all controls included are plotted in Panel (A) and (B) of Figure (V) for scaled applications and approval ratio, respectively. We see that the estimates are mostly similar across all the sub-samples.

Coefficients on  $Treat \times Post$  in Columns (5) and (6) quantify the changes in the growth rate of house prices. The regressions use the census tract-level house price index from [Bogin, Doerner, and Larson \(2016\)](#) that starts in 2000. The coefficients suggest that the growth rate of house prices in the treated areas increased relative to the control by 1.7–1.8 percentage points after

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<sup>16</sup>The average mortgage size in treated tracts in the pre-treatment period was about USD 150,597. Thus the demand for mortgage credit increased by about USD 2.0 million per 1000 adults per census tract ( $USD\ 150,597 \times 13.4$ ), by about USD 5.4 million per treated census tract ( $USD\ 2\ million \times 2.7\ thousand\ adults\ per\ census\ tract$ ), or by about USD 37.8 billion across all treated tracts from bordering counties ( $USD\ 5.4\ million \times 7,011\ treated\ tracts$ ).

the event, though the estimates are statistically significant only at the 10% level. This finding is in line with that of [Di Maggio and Kermani \(2017\)](#): in their sample, house price growth rates increased by 3.3 percentage points following a 10% increase in mortgage origination.

A noticeable limitation of the estimates above is that the mortgage supply in the U.S. had started to shrink from 2005, an antecedent of the financial crisis of 2008, and the post-event regression sample includes the years from 2005 to 2008. Hence, while it may be injudicious to claim that the effects estimated above are completely uncontaminated by these changes, the DID design ameliorates the issue to the extent that the market-wide forces evenly affect the neighboring counties across the state borders. In addition, the effects remain qualitatively and quantitatively similar when estimated in an alternative sample which restricts the post-event period to 2006 (see Section (V), Robustness).

The financial crisis is also often argued to be a result of excessive mortgages taken by borrowers without means, often for investment purposes rather than for occupancy purposes, and this raises the question of whether the increase in the origination reported above is also driven by such borrowers. Table (IV) examines this assertion using the same DID specification. The outcome variable is scaled applications for the owner-occupied mortgages in Columns (1) and (2) and for the non-owner-occupied category in Columns (3) and (4). The coefficients on  $Treat \times Post$  suggest that the applications increased dramatically and significantly only for the owner-occupied category in the treated areas vis-à-vis the control, but not for the latter category. The assertion is further investigated by examining whether there was a compositional change across the two mortgage categories. Columns (5) and (6) examine the scaled applications in the non-owner-occupied category as a fraction of total applications, and Columns (7) and (8) examine the same as a fraction of successful applications. The coefficients in these four columns indicate a modest 1 percentage point increase in non-occupancy mortgages at both the application and origination stage. By and large, the investment-motivated demand does show a slight uptick, but does not appear to be the dominant reason behind the robust 15% increase in mortgage applications.

### (III) Mortgage delinquencies

Results so far imply that more mortgages were originated after credit reports became easier to access. The question then arises, whether mortgage delinquencies too would increase as a result? If the origination increased owing to an improved applicant pool, the delinquencies

would fall, or at least not rise. However, if the origination increased due to subprime lending while the pool stayed the same as before, the delinquencies would rise subsequently.

To examine the patterns in the delinquencies, the GSE data, which are a subset of the HMDA mortgages, are used. First, a mortgage vintage is defined as the collection of the mortgages originated in a given area—treated or control—in a given year—2004 (pre-event) or 2005 (post-event), leading to four vintages: the treated vintage in the pre- and post-event year, and the control vintage in the pre- and post-event year. Then, the rate of delinquency of a given vintage is defined as the ratio of the number of mortgages late on a scheduled payment by  $n$  days for the first time at a given age (measured in months since origination) to the total number of mortgages in that vintage. The rates are analyzed for delays of  $n = 30\text{--}89$  days and  $90\text{--}120$  days.

Panel (A) of Figure (VII) shows the 30–59-day delinquency rate for the treated and control vintages for the year 2004 on the left-hand side and for the year 2005 on the right-hand side. The plot on the left reveals that, among the mortgages originated before the event, the delinquency rates of the treated and control vintages follow almost the same trend; whereas the plot on the right reveals that, among the mortgages originated in the year of the event (2005), the delinquency rate of the treated vintage is *lower* than that of the control vintage. The same pattern is observed for the 60–89-day delinquency rate, plotted in Panel (B) of Figure (VII). Furthermore, each of the delinquency rates of the treated vintage becomes much lower than that of the control vintage during the financial crisis (48 months after 2004, or 36 months after 2005) than during the earlier period.

These reductions may at first appear puzzling, because if lenders use the same screening policy after the event as they were using before the event, one may predict the delinquency rates to stay the same. However, this prediction implicitly assumes that the composition of the pool stays the same, whereas the self-learning mechanism predicts that the composition of the pool may change. New creditworthy borrowers may enter the market and those with poor creditworthiness may drop out. Section (III) provides an evidence consistent with such a change—the proportion of the first-time homebuyers in the originated mortgage pool increased.<sup>17</sup>

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<sup>17</sup> Suppose that lenders deny applicants whose *ex-ante* probability of default falls above some threshold,  $p^*$ . Assume that the average delinquency rate of originated loans in the pre-event period is  $p_1$ , where  $0 \leq p_1 \leq p^*$ . After the free credit report policy is implemented, suppose that an additional pool of applicants is motivated to request a mortgage, and they are subject to the same upper bound,  $p^*$ , but their delinquency rate is  $p_2$ , where  $0 \leq p_2 \leq p^*$ . It is clear that depending on the values of  $p_1$  and  $p_2$ , the average delinquency rate after the event may increase or decrease, given that the delinquency rate of the new entrants is different from that of the older pool.

Overall, the reduction in the delinquency rates after the event suggests that the applicant pool improved, and one reason for this appears to be an increase in the share of first-time home-buyers.

## 4B Mechanism: Consumer Self-learning Channel

This section tests the self-learning mechanism.

### (I) Outcomes in areas likely to have overestimating borrowers

While the self-learning mechanism predicts that the number of mortgage applications after the event should decline for the overestimating borrower type and increase for the underestimating type, empirically testing them is challenging. First, the two borrower types are not distinguishable in the mortgage application or performance data. Second, to the extent that the predictions are true, the mortgage data would not capture the underestimating type, as this type is predicted to *not* apply for credit (mortgage).

Notwithstanding the above limitations, it may be argued that the borrowers from the areas where the *ex-ante* rejection ratios due to DTI ratio were small but due to the credit history were large are more likely to be the over-estimators, relative to the borrowers from other areas. This is because these rejection patterns fit the borrowers who mistakenly overestimate their credit-worthiness, apply for a mortgage, and thus are more likely to be rejected for (bad) credit history than for their repayment inability (high DTI). Following this reasoning, the census tracts in the sample are sorted into *tertiles* of the rejection ratios in the pre-event year 2004 for the DTI ratio and credit history, leading to nine sub-groups.

Table (V) shows the results of separately regressing the scaled applications and approval ratios using Equation (1) for each of the nine sub-groups. The regression results for the scaled applications appear in Columns (1) through (3) and for the approval ratio in Columns (4) through (6). Credit-history tertiles vary from the top to bottom of the table and DTI tertiles, from left to right. The overestimating sub-group corresponds to the third tertile of the credit history and the first tertile of the DTI ratio. We see that the treated areas saw the smallest increase in scaled applications relative to the control areas for the overestimating borrowers (6.23 versus 10.67 or 20.32 within the first DTI tertile) than the other borrowers. This pattern is in line with the prediction for the overestimating borrowers.

The prediction regarding how the approval ratio would change for the overestimating and underestimating type is ambiguous. The ratio should increase for *both* types, but the size of the increase depends on the *ex-ante* relative proportion of the two types in the applicant pool. In Column (4) of Table (V) we see that the amount of increase in the approval ratio is indeed similar across the three credit-history tertiles, though it is not significant for the third tertile.

To summarize, the prediction that after the event the number of applications for the overestimating borrower type should decrease (or should increase less than the others) is largely supported in the data.

## (II) Increase in the mortgage-related cognizance among borrowers

If consumers learn more from their credit reports about their creditworthiness after the event, their decision regarding credit and mortgages will reflect it. In particular, as the reports contain the credit history of consumers, their cognizance of about it should increase. Consumers may thus be able to reduce the likelihood of rejections due to credit history by taking actions such as steps to improve the record before applying for credit or by applying to subprime lenders, who specialize in providing credit to those with poor credit history. At the same time, the likelihood of rejections due to DTI ratio may not change, as it is unlikely that consumers could boost their income strategically before applying for credit.

Similarly, an increase in cognizance would also affect applicants' tendency to withdraw mortgage applications that are still being processed (before the lender has made the decision). It is common for potential applicants to initiate several formal mortgage applications at once at different lenders to hedge against the uncertainty in approvals and mortgage terms. In doing so, they incur multiple non-refundable application costs, but in the end they take out a mortgage with only one lender and withdraw their applications from the others (in-process withdrawals). With an increase in cognizance of their creditworthiness, borrowers' uncertainty over approvals and credit terms decreases, and with that, they are likely to apply to *fewer* lenders at once. Thus the fraction of in-process withdrawals should decrease in the treated areas.<sup>18</sup>

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<sup>18</sup>The withdrawal ratio over the 2000–2008 period is about 12%, indicating that in-process withdrawals are fairly common. Anecdotal evidence suggests that consumers tend to withdraw applications when they find a better offer from other lenders ([Reddit Forum](#), n.d.). More importantly, credit reporting agencies do not penalize multiple applications if they are made within a short time period, as [Equifax](#) (n.d.) explains: “If you’re shopping for a new auto or mortgage loan or a new utility provider, the multiple inquiries are generally counted as one for a given period of time. The length of this period may vary depending on the credit scoring model used, but it’s typically from 14 to 45 days. This allows you to check at different lenders.”



The first prediction is tested by regressing the fraction of total applications rejected for credit history and for DTI ratio. These outcomes are estimated separately for the entire sample and for a sub-sample of only those census tracts where the rate in the pre-event year 2004 was higher than the *regional mean* (Footnote (21) shows the steps to calculate the mean). The reasons to separately focus on the *ex-ante* high-rejection areas are that the information in the reports are more valuable when the rejection rates are high, and the influence of the event on reason-specific rejection probabilities will be greater in the areas where rejections were frequent before the event. The second prediction is tested by regressing the withdrawal ratio, which is the fraction of total applications that are formally withdrawn by borrowers before lenders could make a decision.

These predictions are tested using the regression specification from Equation (1), and Table (VI) shows the results. In Columns (1) through (4) we see that the fraction of applications denied due to credit history decreased by 0.3 percentage points in the treated tracts relative to the control, statistically significant in the *ex-ante* high-rejection-rate areas (Columns (3) and (4)). The coefficients in columns (5) through (8) show that the DTI ratio denials did not decrease statistically significantly. Though the estimates carry only modest statistical significance, they indicate that the reason-specific rejection likelihoods changed in a manner consistent with potential borrowers becoming more cognizant of their credit history.<sup>19</sup> The estimates for the withdrawal ratio appear in Columns (9) and (10) and imply that it decreased by 0.9–0.11 percentage points in the treated tracts vis-à-vis the control.<sup>20</sup>

Overall, these findings point to an increase in mortgage-related cognizance among borrowers, consistent with the self-learning mechanism.

#### 4C Characterizing the Effect: Who benefits?

Characterization of the consumers and the areas that are more likely to benefit from easier access to credit reports may provide insights about those for whom the information frictions about creditworthiness are likely to be binding, and it may also be useful for policy targeting.

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<sup>19</sup> A caveat of this analysis is that HMDA does not mandate lenders to report reasons for rejections, so if the reporting incentives of lenders were also influenced by the event, the estimates reported above would be the result of the changes in borrowers' cognizance and lenders' incentives. However, the incentives to report rejection reasons would need to change in the event year in a particular manner that varies across the treated and control areas, even for lenders that may operate in both areas. Such precise changes in incentives for reporting the reasons for rejections appear unlikely. Moreover, lenders reported reasons for rejection in 70.81% of the rejected applications in the sample.

<sup>20</sup> In economic terms, the drop is equivalent to ~2.34 fewer in-process withdrawals per treated tract or ~16,513 fewer withdrawn applications aggregated over the treated border counties. At an average cost of ~USD 400 per withdrawn application, this represents ~USD 6.6 million saving in upfront mortgage application fees.

The heterogeneity in the treatment effect across consumer creditworthiness and income is examined next.

### (I) High creditworthy areas

Given that creditworthy borrowers are more likely to be granted mortgage credit, if credit reports aid consumers in assessing creditworthiness, easier access to the reports should lead to a *greater* increase in applications and approvals in areas where the fraction of creditworthy consumers was higher before the event. This is because, under the self-learning mechanism, the drop out would be larger and the entry smaller in the subprime areas vis-à-vis the prime areas.

To test this, a county is classified as having high creditworthiness if its subprime population fraction is less than the *regional mean* before the event.<sup>21</sup> The year 1999 is chosen as the classification year because [Mian and Sufi \(2009\)](#) suggest that such classification should be done at a time well before the start of a housing boom, as creditworthiness of an area may endogenously evolve with the boom. The earliest year for which data on the county subprime fraction is publicly available is 1999 ([FRBNY & Equifax, n.d.](#)).

Table (VII) shows the results of regressing scaled applications and approval ratio separately using regression Equation (1) for counties with high and low creditworthiness. Within the *ex-ante* high-creditworthiness counties, vis-à-vis the control counties, the treated counties saw an increase of 16.8–18.8 (17.4–19.5%) in scaled applications (Columns (1) and (2)) and a 2 percentage points increase in approval ratio (Columns (3) and (4)). At the same time, within the *ex-ante* low-creditworthiness counties, vis-à-vis the control counties, the treated counties saw an increase of just 8.59–11.66 (8.8%–12%) in scaled applications (Columns (5) and (6)) and a 1 percentage point increase in approval ratio (Columns (7) and (8)), which is statistically significant only at 10%. Taken together, these estimates support the self-learning mechanism and suggest that creditworthy borrowers are more likely to benefit from easier access to credit reports.

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<sup>21</sup> The steps to calculate *regional mean* are as follows. A region is defined as the area encompassing a control (pre-FACTA) state and all the surrounding states. Consider the control state Colorado (CO) and all the surrounding treatment states. The regional mean for this region is the average rejection rate for the census tracts in all the counties at the border between CO and WY, UT, AZ, NM, OK, KS and NE. The regional means of rejection rates for all seven control states are calculated in this way, and a census tract is then classified as a “high rejection tract” if its rejection rate is more than the regional mean in 2004.

## (II) Low income borrowers

The treatment effect may vary across borrowers of different incomes, because the consequences of a mortgage rejection are more severe for low-income borrowers. Thus, upon learning one's creditworthiness, the likelihood of not applying for credit (drop out from the pool) or gravitating to subprime lenders is higher for low-income borrowers as opposed to high-income borrowers. Alternatively, the over- and under-estimation tendencies under the self-learning mechanism also lead to the same prediction. As lower income is associated with a higher likelihood to overestimate one's creditworthiness ([Perry, 2008](#), Table III), low-income consumers are more likely to revise their creditworthiness downwards after learning their true creditworthiness from their credit reports, and thus are more likely to avoid rejection costs by not applying for credit. The approval ratio for this sub-group of consumers is thus likely to increase owing to the drop out channel of the self-learning mechanism, whereas it is unclear how the ratio would change for high-income consumers.

To test these predictions, first, the cut-offs for the income quartiles are calculated each year within the sample, and then the applications from each quartile are aggregated to the census-tract level and scaled by the population (measured in 1000's). The approval ratio is then calculated within each quartile.

Panel (A) of Table ([VIII](#)) shows the results of regressing the scaled applications separately for each of the income quartiles using Equation ([1](#)). We see that the scaled applications did not increase significantly for the lowest quartile, but increased significantly for the other three, and the increase was larger and statistically significant for these quartiles. The significant increase in applications among the higher quartiles but not among the lowest-quartile consumers is consistent with the prediction that it is the latter for whom drop out is more likely.

Panel (B) of Table ([VIII](#)) shows the results of regressing approval ratios separately for each of the income quartiles using Equation ([1](#)). We see that the ratios increased statistically significantly in the treated areas relative to the control areas only for the lowest income quartile borrowers, consistent with the prediction. The high-income borrowers did not see an increase in approvals likely because the marginal propensity of lending to such consumers is larger than that of lending to low-income consumers ([Agarwal, Chomsisengphet, Mahoney, & Stroebel, 2018](#)), thereby such consumers are not the ones who are excluded from the credit markets. The absence also re-emphasizes that the value of accessing credit reports is relatively higher for the low-income borrowers.

### (III) First-time homebuyers

First-time homebuyers tend to be younger adults. The 25<sup>th</sup> percentile and median of their age is 29 and 35 years, respectively (Raymond & Dill, May 20, 2015). Younger adults account for disproportionately high fraction of those lacking robust credit records (Federal Reserve Board of Governors, 2007, p.28). Hence, proportion of first-time homebuyers may increase after access to credit reports becomes easier.

The prediction for the first-time homebuyers is tested next. The outcome variable is defined as the ratio of the number of mortgages taken out by first-time homebuyers to the number of all originated mortgages that had known information on first-time homebuyer status. It is important to enumerate two limitations of this analysis. Whether an applicant is a first-time homebuyer is recorded only in the GSE data, not in the HMDA data. Also, since the property location information in the GSE data is only available at the 3-digit zip code level, the properties were mapped to the counties using simplified approximations (see Footnote (28) in Data Appendix). Accordingly, the regression is specified at the zip3-state level as follows:

$$Y_{zsjt} = \beta_0 + \beta_1 \times \text{Treat}_{zsj} \times \text{Post}_t + \delta \times \text{Economic controls} + \alpha_{zs} + \gamma_{jt} + \varepsilon_{zsjt} . \quad (3)$$

Here,  $z$  indexes the areas delineated by a 3-digit zip code at the border of treated state  $s$  and control state  $j$ .  $\alpha_{zs}$  is zip3-state fixed effects.  $\gamma_{jt}$  is the *Border*  $\times$  *Quarter* fixed effects, and it serves similar function as that of the *Border*  $\times$  *Year* fixed effects in Equation (1). The sample is limited to the zip3-state areas that come under the border counties of treated and control states.

Columns (1) and (2) of Table (IX) show the regression results. The coefficients suggest that the percentage of first-time homebuyers increased by 1 percentage point in the treatment areas relative to the control areas.<sup>22</sup> This finding is in line with the prediction that, following the event, more new entries should occur in the treated areas.<sup>23</sup>

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<sup>22</sup> About 6.7% of the observations within the homebuyer data sample pertaining to the bordering counties do not have information on first-time homebuyer status. The specifications that alternatively define the outcome variable as the ratio of number of first-time homebuyers to *all mortgages* yield similar estimates, and these estimates are left unreported for brevity.

<sup>23</sup> A concern with this estimation is that the mortgage sample consists of only those that were purchased by the GSEs. However, as argued before, this selection would be an issue only if GSEs' incentives to purchase first-time homebuyer mortgages relative to their overall purchase from the treated counties increased relative to the control counties from the event year 2005 onward. Such a time- and location-specific change seems improbable.

#### 4D A Demand- or Supply-side Effect?

The findings so far indicate that the likely explanation for the increase in origination in the treated areas relative to the control after the event is the self-learning mechanism, which is a demand-driven channel. However, a supply-driven explanation is also plausible. Even though it was only the consumers for whom credit reports became easier to access, lenders too understood this and could have responded by increasing the mortgage supply.

Even though the supply-driven explanation is plausible, many of the earlier findings favor the demand-driven explanation. First, an increase in applications and a decrease in in-process-applications withdrawals are a result of decisions that are taken solely by potential borrowers, and these quantities are mostly independent of lenders' influence. Second, it is the demand-driven mechanism under which the effects would be heterogeneous across consumer characteristics, as it was across creditworthiness and income. Additionally, given that the propensity of lenders to extend credit to low-income borrowers is low (Agarwal et al., 2018), and given that in the current setting we see that the approval ratio increased significantly for such borrowers vis-à-vis the high-income borrowers, the supply-driven explanation appears unlikely.

Notwithstanding the above suggestive evidence favoring the demand-side explanation, two outcomes more directly related to supply-side characteristics are examined next: mortgage interest rates and heterogeneous effect by *ex-ante* density of mortgage lenders.

##### (I) Mortgage interest rates on the GSE-repurchased mortgages

The changes in the mortgage interest rates after the event in treated and control areas can be utilized to examine whether the increase in origination was supply- or demand-driven. If it is the former, the rates would decrease; if the latter, they would increase.

The investigation of the rates needs to account for characteristics of the property and borrower-risk. Scharfstein and Sunderam (2016) argue that the prices (interest rate) at which lenders sell *conforming* mortgages to the GSEs materially vary only across three dimensions: credit score, loan-to-value ratio (LTV), and loan type (adjustable rate, fixed rate etc.).<sup>24</sup> Therefore, the residuals obtained from a regression of the rate on these dimensions approximately measure the lender-specific pricing schedule independent of the characteristics of borrowers and the mortgage. In the context of the current paper, only the first two

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<sup>24</sup>Fannie Mae's mortgage pricing variation across these dimensions can be seen in its pricing schedule here: <https://singlefamily.fanniemae.com/media/document/pdf/llpa-matrix-pdf>

dimensions are relevant, since the GSE sample only includes one type of mortgage—the 30-year fixed-rate single-family mortgage. Thus, when the rate is regressed on credit score, LTV, and the  $Treat \times Post$  interaction term, the coefficient on the last term captures the change in the pricing schedule of lenders in the treated areas vis-à-vis the control areas after the event. If the lenders lowered the mortgage interest rates in the treated areas (in a bid to increase mortgage origination), the sign on the coefficient would be negative; if they raised the rates, the sign would be positive.

The regression specification is similar to Equation (3), but it is now specified at the loan level  $i$  as follows:

$$\text{Interest Rate}_{izsjt} = \beta_0 + \beta_1 \times \text{Treat}_{izsj} \times \text{Post}_t + \delta \times \text{Controls} + \alpha_{zs} + \gamma_{jt} + \varepsilon_{izsjt} . \quad (4)$$

Columns (3) and (4) of Table (IX) show the results of the regression. In Column (3), *Controls* are the two relevant pricing dimensions, credit score and CLTV (combined loan-to-value: the loan-to-value ratio inclusive of all loans secured by a mortgaged property); and in Column (4), to make the specification more rigorous, *Controls* additionally include DTI ratio, number of units comprising the mortgaged property, and percentage of mortgage insurance. Results show that the coefficients on  $Treat \times Post$  in both the columns are positive and significant, at about 0.009–0.01 percentage points. Thus, if anything, lenders responded to the event by increasing, not decreasing, the risk-adjusted mortgage interest rates in the treated areas relative to the control, contrary to what a supply-driven explanation would predict.<sup>25</sup>

## (II) Heterogeneous effects by *ex-ante* density of mortgage lenders

If the increase in mortgage origination were driven by lenders, it would be greater in areas where the *ex-ante* density of lenders is high. To examine this, first, census tracts are classified into high and low lender-density groups: high if the number of HMDA mortgage lenders per adult in the pre-event year 2004 in a census tract was greater than the *regional mean* (defined in Footnote (21)), and low otherwise.

Columns (1) through (4) of Table (X) show the results of separately regressing, using Equation (1), the dollar origination volume (in 1000 USD) per adult for the two density groups. The

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<sup>25</sup>The magnitude of the increase in the rates is tiny, potentially for two reasons. First, interest rates on conforming (GSE-repurchased) loans do not vary across regions or with dimensions other than FICO scores, loan-to-value ratio, and loan type (Hurst, Keys, Seru, & Vavra, 2016). Second, the supply of mortgages in the U.S. is highly elastic because of the large-scale purchases by the GSEs of conforming mortgages in the secondary market.

estimates are smaller in magnitude and have weaker statistical significance for high-density tracts (Columns (2) and (4)) vis-à-vis the respective low-density census tracts (Columns (1) and (3), respectively). Thus, among the *ex-ante* high-lender-density census tracts, relative to the control tracts, the treated tracts saw a much *smaller* increase than the treated tracts from the low-lender-density group relative to the respective control tracts. Moreover, the t-test for the difference in the coefficient of  $Treat \times Post$  in high- and low- lender-density areas (*High – Low*) shows no statistical difference.

Columns (4) through (8) of the table show the above regressions for the approval ratio. The results are similar—there is no statistical difference in the increase in the approval ratio in areas with a high or low lender density.

The finding that the effects were stronger, not in the areas that had *ex-ante* high lender density, but in the areas that had an *ex-ante* low lender density is inconsistent with the supply-driven explanation.

## 5 Supplementary Discussion

### (I) Did origination increase due to rise in private mortgage securitization?

As private securitization of mortgages (selling the mortgage to non-government agencies) offers lenders higher commissions, a rise in the tendency to privately securitize could explain the increase in the mortgage origination of reported in the current paper ([Keys, Mukherjee, Seru, & Vig, 2010](#)). If this explanation is true, the fraction of originated mortgages that were sold to non-government entities would increase in the treated areas relative to the control after the event.

Table (XI) shows the result of regressing, using Equation (1), three outcome variables: the fraction of total applications that lenders originated and (i) sold to non-government entities, (ii) sold to the four GSEs (Fannie Mae, Freddie Mac, Ginnie Mae, and Farmer Mac), and (iii) did not sell. The estimates suggest that private securitization did not increase (Columns (1) and (2)); government securitization increased (Columns (3) and (4)); and the fraction of unsold mortgages did not change (Columns (5) and (6)). Overall, private securitization does not appear to a reason behind the increase in origination.



## (II) Did origination increase due to subprime lending? Credit score-based evidence

It may be argued that the increase in mortgage origination was due to an increase in the subprime credit (Mian & Sufi, 2009). Using the comprehensive HMDA data and location-based proxies of creditworthiness, Table (VII) already suggests that the effect of free credit reports was stronger in the prime counties/census tracts than in the subprime. Even though these proxies are informative and widely used (e.g., Di Maggio & Kermani, 2017; Mian & Sufi, 2009), they are imprecise. The GSE sample contains the application-level credit scores, and thus it can be used to precisely examine the patterns in prime and subprime origination.

Table (XII) shows the results of regressing separately the number of prime (credit score  $\geq 620$ ) and subprime *originated* mortgages in zip3-state areas using Equation (3). Columns (1) and (2) show that the number of prime mortgages increased by 308–312 in the treated zip3-state areas relative to similar control zip3-state areas, whereas columns (3) and (4) imply that subprime mortgages increased only by ~10 applications, which is 30 times smaller. Thus the increase in mortgages did not disproportionately go to subprime consumers. Note that these estimates are not directly comparable to the previous regressions, as the observation unit here is zip3-state, not census tracts, and the outcome variable is not scaled by population. Also, these results suffer from the same selection issue that existed with the previous results utilizing the GSE sample, and the same argument that was made before—the implausibility that the incentives of GSEs changed across the contiguous sample counties around the event—alleviates it. In addition, Elul, Gupta, and Musto (2020) show that to combat the onset of the housing bust before 2007, the GSEs sought to buy more subprime, not prime, mortgages. Hence, had the GSEs not changed their buying pattern, the estimates of subprime origination would have been even lower and the contradiction of the subprime hypothesis even stronger.

## (III) Effect on banks

The analysis in the paper so far has focused on evaluating the effects on borrowers, but it is the lenders who ultimately evaluate the credit decisions, and hence the effect on banks is evaluated next.

Analyzing the effects on banks has a few notable limitations. First, banks are not the dominant mortgage originators. Despite being 80% of mortgage lenders by number, banks accounted for just 37% of mortgage lending in 2005 (Avery, Brevoort, & Canner, 2007), and thus the conclusions drawn from studying banks may not generalize to all mortgage lenders. Second, since

many banks operate across states, their treatment and control status in this natural experiment is not binary, but continuous. To solve this issue, it is assumed that the continuous treatment intensity is proportional to a bank's *ex-ante* mortgage origination in the treated states relative to the activity in treated and control states combined. Following this, a bank is classified as "control" if in the pre-event year 2004 the ratio of mortgage amounts it originated in the control states to those in the treated and control states combined was larger than the cross-sectional mean in that year across all banks in the sample. The regression equation is

$$Y_{bt} = \beta_0 + \beta_1 \times \text{Treat}_b \times \text{Post}_t + \delta \times \text{Bank controls}_{bt} + \alpha_l + \gamma_t + \varepsilon_{bt} , \quad (5)$$

where  $Y_{bt}$  represents the three outcome variables: net interest margin (NIM), return on equity (RoE), and return on assets (RoA);  $b$  indexes the banks;  $\text{Treat}_b$  is 1 if a bank is treated and 0 otherwise;  $\text{Post}_t$  is 1 if  $\text{year} \geq 2005$  and 0 otherwise; year  $t$  represents year-quarter;  $\alpha_l$  is bank fixed effects;  $\gamma_t$  is year-quarter fixed effect; and *Bank controls* include banks' log total assets, share of liquid assets to total assets, and cost of deposit.<sup>26</sup>

The regression results in Table (XIII) show that treated banks saw a 6 basis-points increase in NIM (Columns (1) and (2)), a 0.75–0.76 percentage-points increase in RoE (Columns (3) and (4)), and a 0.07–0.08 percentage-points increase in RoA (Columns (5) and (6)). Also, the results are qualitatively and quantitatively similar when lenders are classified into treatment and control groups using cross-sectional median instead of the mean.

Furthermore, it is also useful to understand which characteristics of the banks allowed them to increase the mortgage supply after the event. To this end, the role of liquid assets on a bank balance sheet is examined. A bank is classified as having a *high* share of liquid assets if its ratio of liquid assets to total assets (computed annually) was *greater* than the cross-sectional mean in the pre-event year 2004, and as having a *low* share of liquid assets otherwise. The total amount of mortgages (in million USD) originated by a bank in the control and treated states is regressed using Equation (5) separately for these two groups. Also, since the mortgage information come from the HMDA data, which are publicly available only at an annual frequency, the regression is estimated at the *Bank*  $\times$  *Year* level.

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<sup>26</sup> NIM is the ratio of net interest income (sum of RIAD4074 and RIAD4301) to earning assets. I use the definition of earning assets from St. Louis Fed: it is the sum of RCFD0071, RCFD1350, RCFD2122, RCFD3545, RCFD1754, and RCFD1772 (<https://fred.stlouisfed.org/series/USNIM>). RoE is the ratio of net income (RIAD4340) to book value of equity. RoA is the ratio of net income to book value of total assets. Liquid assets is the sum of RCFD1754, RCFD1773, RCFD3545, RCFD1754, RCFD3545, and RCFD1350. Cost of deposit is the ratio of RIAD4073 to earning assets.

Panel (B) of Table (XIII) shows the results of the regressions. The coefficients in Columns (1) and (2) suggest that, among the banks with an *ex-ante* low share of liquid assets, the change in mortgage lending by treated banks vis-à-vis the control banks is not statistically significant, and the point estimate is negative. At the same time, the coefficients in Columns (3) and (4) suggest that among the banks with an *ex-ante* high share of liquid assets, the treated banks on average increased mortgage lending relative to the control by about USD 111 million, in a statistically significant manner.

All in all, the effect of the event on financial performance of the banks seems to be positive, and the banks with high liquid assets appear to be behind the increase in mortgage origination after the event.

#### (IV) An alternative mechanism based on information asymmetry

An alternative mechanism based on asymmetric information in which borrowers *privately know* their true creditworthiness type, but do not know what lenders know about them, is plausible. Using free credit reports, borrowers learn that the creditworthiness information on them available at lenders are proportional to their true creditworthiness type. Hence, under the non-trivial search/application cost, borrowers with poor creditworthiness (bad type) self-select out. The applicant pool now improves relative to the situation in which borrowers do not know that a lender has information about their true type and optimistically expect that the information is better than what is warranted by their credit reports. Note that the improvement occurs here due to the self-selecting-out by bad type, but not by self-selecting-in by good type, since all borrowers *privately know* their true type. However, since under the self-learning mechanism, borrowers themselves have imperfect information of their true type, both selecting-in by good borrowers and selecting-out by bad borrowers contribute to pool improvement after credit reports become free.

The empirical findings are consistent primarily with the self-learning mechanism. We saw that in the treated areas both the mortgage applications and first-time homebuyers fraction increased, not decreased. Both these findings provide evidence of selecting-in by borrowers, which is plausible only under the *self-learning* mechanism.

Another valid concern is that in assessing mortgage applications, together with the credit reports, lenders use private information that they may accumulate through relationship lending.

This attenuates the effects of free credit reports. The concern is partially alleviated by the fact that lenders necessarily look at credit reports and scores when assessing credit applications.<sup>27</sup>

## (V) Robustness

Since the natural experiment utilized in this paper occurred in the year 2005, the sample period is chosen to be from 2000 to 2008 to allow for enough post-experiment observations. As the experiment is close to the financial crisis of 2008, it is crucial to ensure that the results are not caused by the unique lending environment that existed in 2007–2008. To this end, all the regressions were re-estimated by excluding the observations for the years 2007 and 2008. Mostly the results are qualitatively and quantitatively similar and are left unreported for brevity.

## 6 Conclusion

Several large-scale surveys in the U.S. indicate that a non-trivial proportion of consumers do not check their credit reports and do not know their credit scores. At the same time, data from retail mortgage market reveal patterns that suggest that many consumers misjudge their creditworthiness when making mortgage decisions. Motivated with the idea that credit reports may aid consumers to self-assess their creditworthiness, this paper uses a natural experiment to examine how do the mortgage market outcomes change when accessing their credit reports becomes easier for consumers.

The federal *Fair and Accurate Transactions Act of 2003* (FACTA) has made access to credit reports easy and free through a website since 2005 for all U.S. consumers. It took just a few clicks to access one's reports at the website, whereas earlier one needed to make a request for a report by calling or mailing a letter to the credit reporting agencies. Even after successfully making the request, one often needed to wait more than a week to receive the reports through mail. Prior to FACTA, seven states had local laws that made it easy for state's residents to access their credit reports and drove the usage of credit reports locally many times higher relative to the usage in the rest of the states.

This paper exploits the pre-existing differences in consumer usage of credit reports in the states where the reports were free versus where it was not in a difference-in-differences setting.

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<sup>27</sup> Experian (n.d.) explains: "Not all lenders think the same way, and they may have different ways of making their decisions. But all of them will look at some key factors to help them decide. These include: information on your credit report including your credit history and public record data."

The counties at the border between the early-adopting states and the contiguous states constitute the sample, former serving as the control group and the latter as the treatment group. The primary finding is that the applicant pool in the mortgage market improved after it became easier for consumers to access their credit reports, as approval ratios increased and subsequent delinquencies decreased. Moreover, mortgage applications increased, more credit was originated to creditworthy borrowers, more first-time homebuyers took out mortgages, and financial performance of mortgage-lending banks improved.

Taken together, the findings support the idea that easier access to credit reports to consumers aids them in making better mortgage-related decisions. Any policy intervention aimed at educating consumers of their creditworthiness may bring about similar improvements in other retail credit markets as well.

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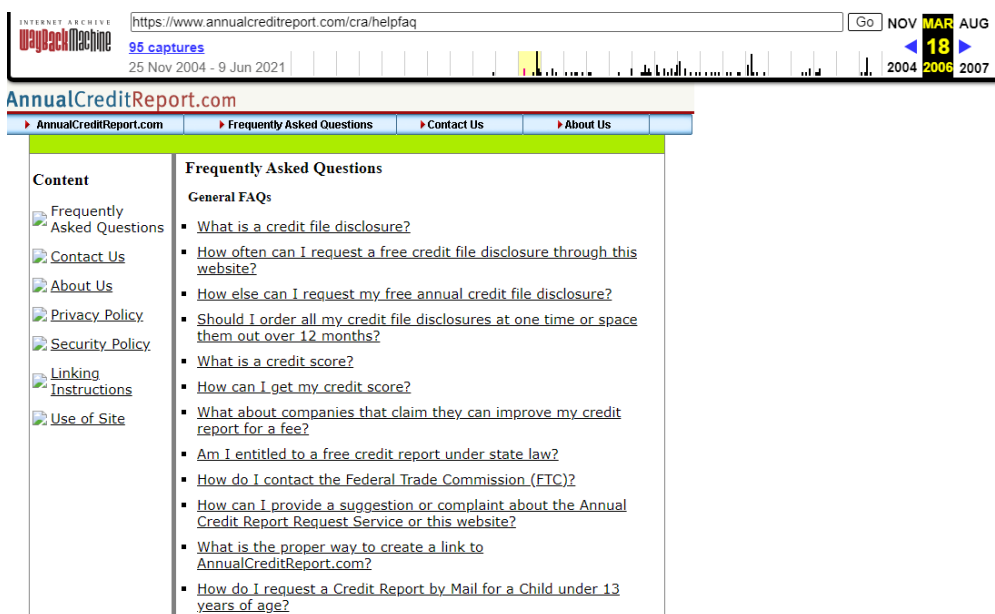
Figure I: The Website

Panel (A) of this figure shows a screenshot of the homepage of the website [www.annualcreditreport.com](https://www.annualcreditreport.com), as it existed on September 25, 2005. Panel (B) of this figure shows a screenshot of the Frequently Asked Questions section of the website, as it existed on March 18, 2006. Both the screenshots are taken from the archive of the website stored at the Wayback Machine.

Panel A: Homepage of the Website



Panel B: Frequently Asked Questions Section of the Website



**Figure II: A Sample Credit Report**

This figure shows the summary page of a free credit report obtained from the website [www.annualcreditreport.com](http://www.annualcreditreport.com) for free under the *Fair and Accurate Transaction Act of 2003*. The specific credit history-related details are not shown. The report contains, among other things, the details of the consumer's active accounts, debt-to-credit ratio, and an indication of the available borrowing capacity.

## 1. Summary

Review this summary for a quick view of key information contained in your Equifax Credit Report.

<b>Report Date</b>	Apr 14, 2020
<b>Credit File Status</b>	No fraud indicator on file
<b>Alert Contacts</b>	0 Records Found
<b>Average Account Age</b>	5 Months
<b>Length of Credit History</b>	8 Months
<b>Accounts with Negative Information</b>	0
<b>Oldest Account</b>	DISCOVER BANK (Opened Aug 29, 2019)
<b>Most Recent Account</b>	AMERICAN EXPRESS (Opened Jan 10, 2020)

### Credit Accounts

Your credit report includes information about activity on your credit accounts that may affect your credit score and rating.

Account Type	Open	With Balance	Total Balance	Available	Credit Limit	Debt-to-Credit	Payment
Revolving	2	2	\$606	\$11,044	\$11,650	5.0%	\$70
Mortgage							
Installment							
Other							
Total	2	2	\$606	\$11,044	\$11,650	5.0%	\$70

### Other Items

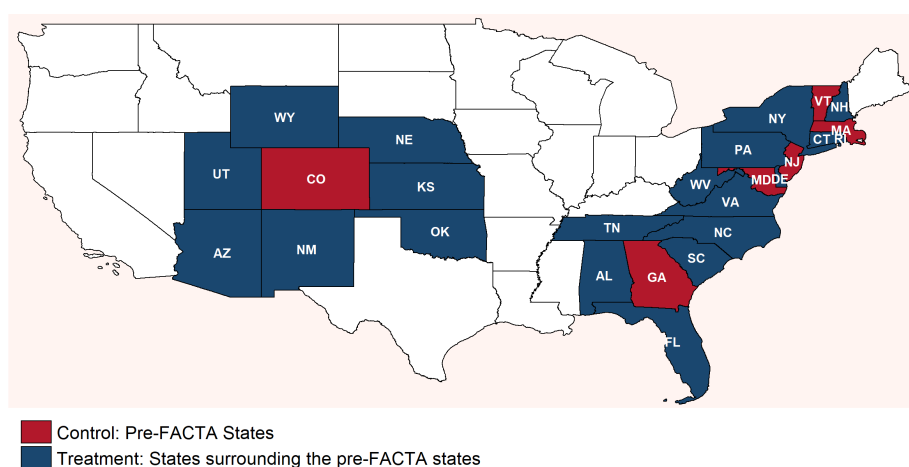
Your credit report includes your Personal Information and, if applicable, Consumer Statements, and could include other items that may affect your credit score and rating.

<b>Consumer Statements</b>	0 Statements Found
<b>Personal Information</b>	3 Items Found
<b>Inquiries</b>	2 Inquiries Found
<b>Most Recent Inquiry</b>	DISCOVER BANK Aug 27, 2019
<b>Public Records</b>	0 Records Found
<b>Collections</b>	0 Collections Found

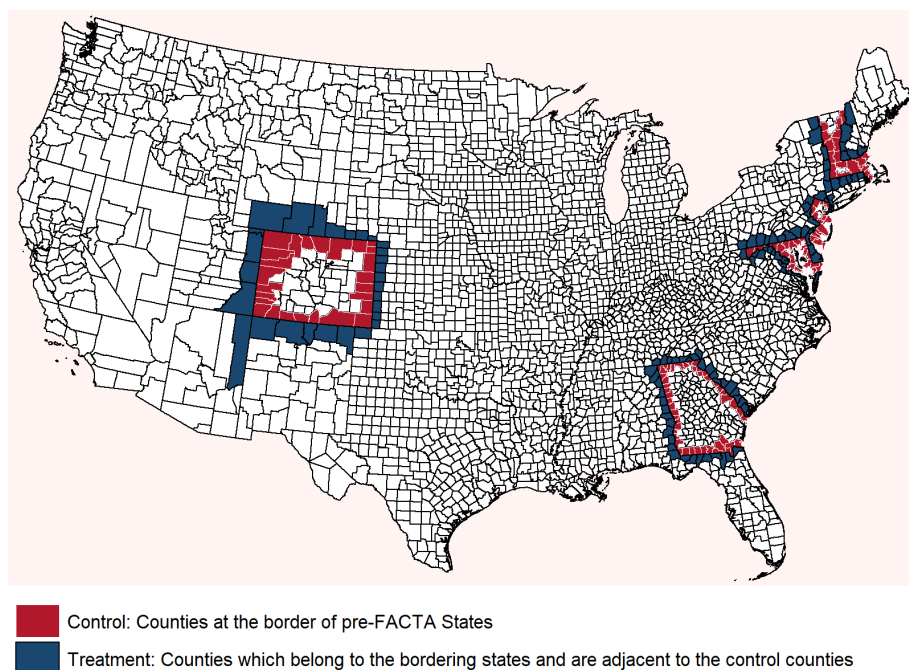
**Figure III: Empirical Research Design**

**Panel (A)** of this figure shows on the map of the contiguous U.S. the states utilized in the difference-in-differences (DID) setting. Seven U.S. states had enacted free credit report laws prior to the FACTA enactment in 2004: CO (1997), GA (1996), MD (1992), NJ (1997), MA (1995), VT (1992), and ME (2003). All except ME constitute the control group, and the 26 states surrounding the control group are the treatment. **Panel (B)** of this figure shows on the map of the contiguous U.S. the counties included in the estimation sample. These are the counties at the border between the treatment and control states.

**Panel A: Treatment and Control States**



**Panel B: Sample Counties from the Treatment and Control States**



**Figure IV: Examining the Parallel Trends**

**Panel (A)** of this figure shows the mean approval ratio in the treated and control census tracts.

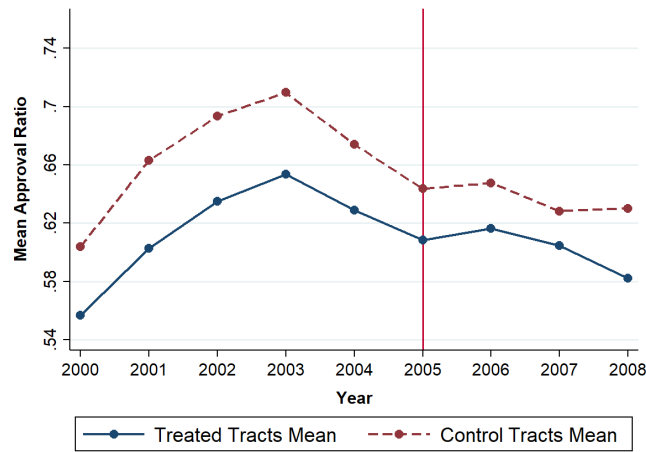
**Panel (B)** of this figure shows the coefficients  $\beta_k$  from regressing *Approval Ratio* using the following specification:

$$Y_{icsjt} = \beta_0 + \sum_{k=T-3}^{T-1} \beta_k \text{Treatment}_{icsj} \times \text{Event}_k + \sum_{k=T+1}^{T+4} \beta_k \text{Treatment}_{icsj} \times \text{Event}_k + \alpha_i + \gamma_{j,t} + \varepsilon_{icsjt},$$

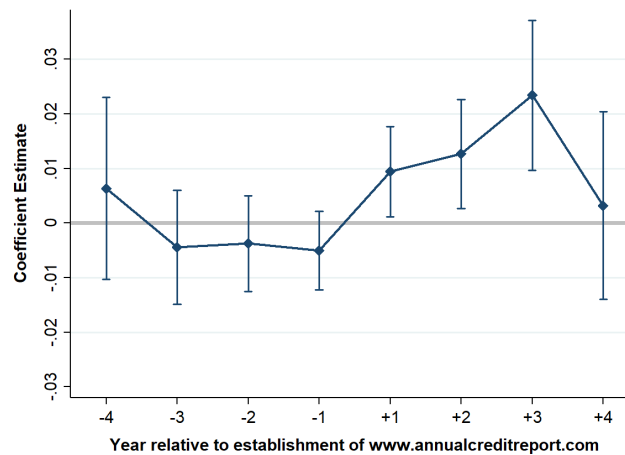
where  $\text{Event}_k = 1$  if  $t = T - k$ .  $\text{Event}_k = 0$  if  $t \neq T - k, k = \{-3, 4\}$ .  $T = \text{Event year 2005}$ .

Coefficients are estimated with respect to the base year 2004 ( $j = 0$ ). The  $x$ -axis shows year relative to the pre-event year 2004; i.e.,  $T = +1$  is the first treated year, 2005. The  $y$ -axis shows the coefficients  $\beta_k$ . The 95% confidence intervals of  $\beta_k$ s are also shown. The regression includes *Border*  $\times$  *Year* and *Census Tract* fixed effects. Other terms in the equation are the same as those in Equation 1. Standard errors are clustered by county.

**Panel A: Mean Approval Ratio in Treated and Control Areas**



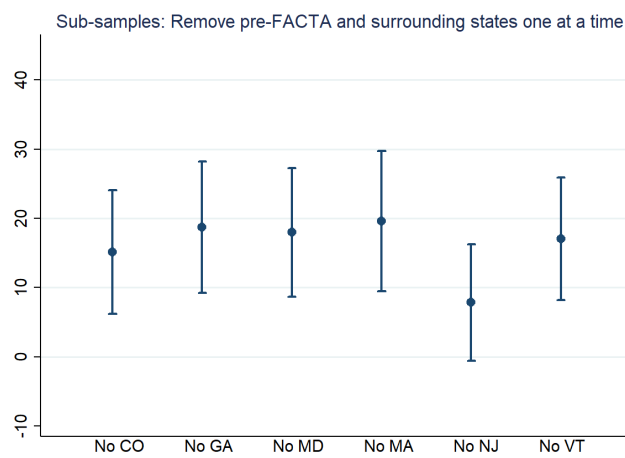
**Panel B: Coefficient Estimates of Approval Ratio by Years to Treatment**



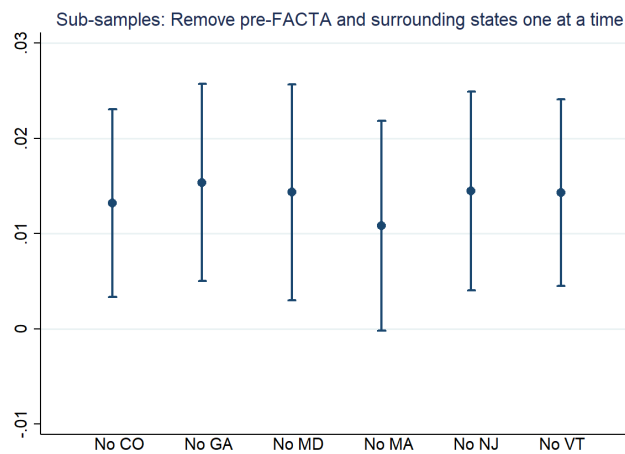
### Figure V: Subsample Analysis

**Panel (A)** of this figure shows the estimates for changes in number of applications per 1000 adults (scaled applications) when each control state is removed one by one. **Panel (B)** of this figure shows the estimates for changes in approval ratio when each control state is removed one by one. For example, the coefficient corresponding to “No CO” represents the estimate when Colorado and its surrounding states are removed from the estimation sample. The regressions specifications behind the estimates are the same as those in Table(III). The bands around the estimates show 95% confidence intervals.

**Panel A: Number of Applications (per 1000 Adults in a Census Tract)**



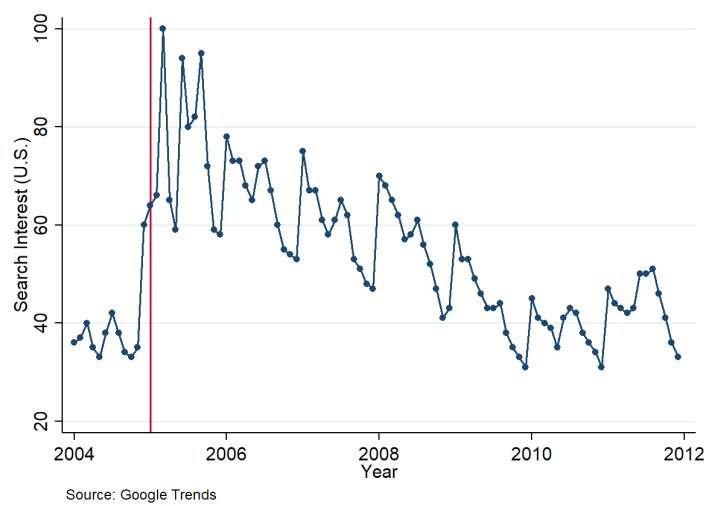
**Panel B: Approval Ratio**



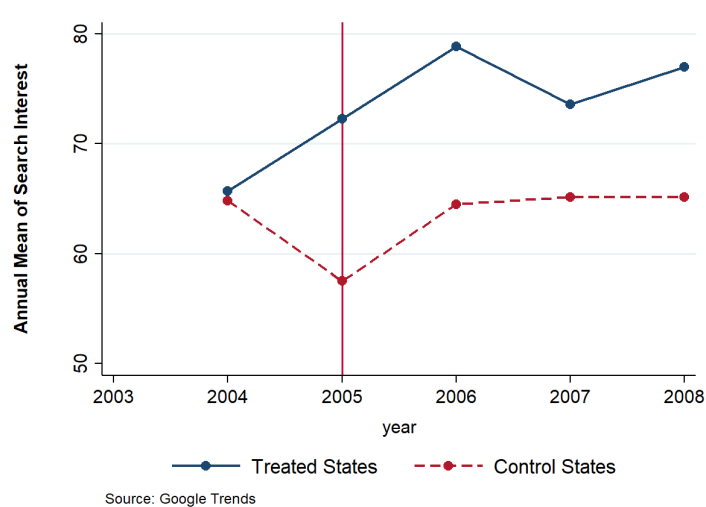
**Figure VI: Search Interest in Free Credit Reports: Google Trends**

This figure plots the search interest in free credit reports using Google Trends data. **Panel (A)** of this figure shows the plot of *Search Interest* for the keyphrase *Free Credit Report* in the US from Jan 1, 2004 till Dec 31, 2011. Numbers on the vertical axis represent search interest relative to the highest point on the chart during this period. A value of 100 (50) represents the peak popularity (half of the peak popularity) for the keyphrase. A value of 0 means there was not enough data. **Panel (B)** of this figure shows the difference in the mean search interest for treated and control states for the same keyphrase from 2004 to 2008 using the interest-by-subregion data from Google. These data are computed within the time period for which the data are extracted from Google. A value of 100 represents the location with the highest popularity of the search term as a fraction of total searches in that location, and a value of 50 indicates a location where it is half as popular. To overcome the issue of data-value normalization by Google, first, the data were *separately* extracted for each one-year interval, and then the means were *separately* calculated within each time interval for each of the two sets of states.

**Panel A: Time Series Search Interest in the U.S. for the Terms “Free Credit Report”**



**Panel B: Cross-sectional Search Interest for the Terms “Free Credit Report”**

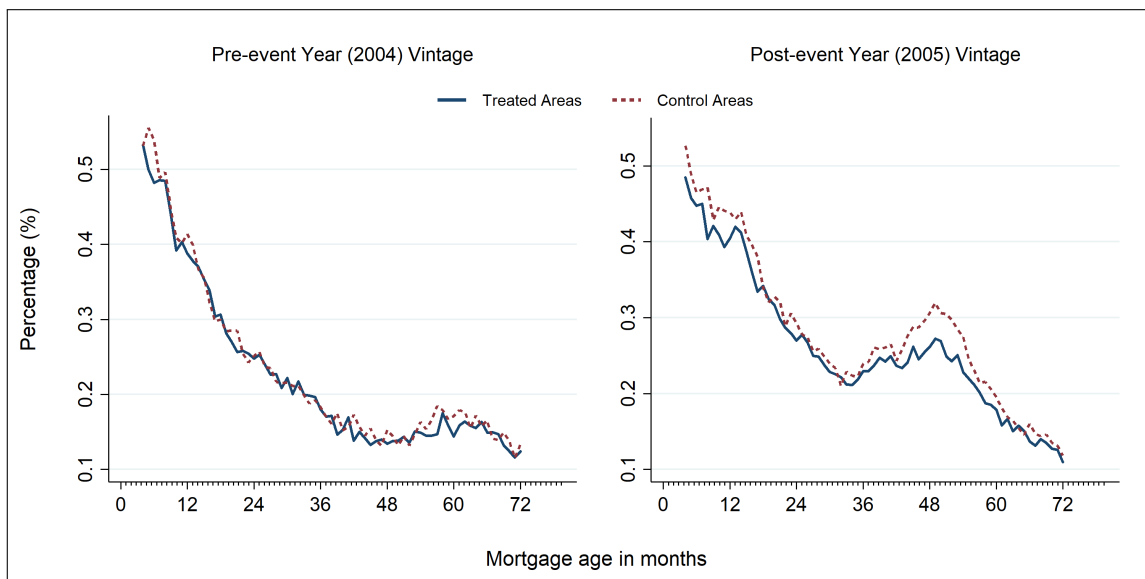




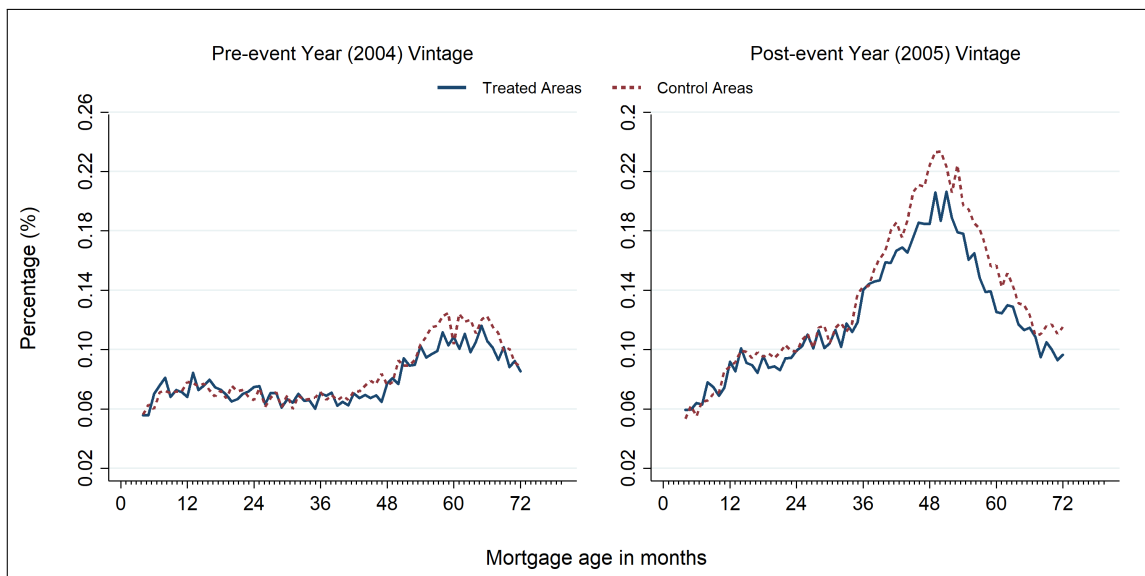
**Figure VII: Effect of Free Credit Report on Mortgage Delinquencies**

This figure plots the delinquency rates of various mortgage vintages with their age (measured in months). A mortgage vintage is a collection of mortgages originated in a given area—treated or control—in a given year—2004 (pre-event) or 2005 (post-event). Delinquency rate is calculated at each age as the ratio of the number of mortgages becoming delinquent for the first-time to the total number of mortgages in the respective vintage. **Panel (A)** shows 30–59-day delinquency rates separately for treated and control areas for 2004 on the left-hand side and for 2005 on the right-hand side. **Panel (B)** shows these same plots for 60–89-day delinquencies. These plots are based on the 30-year fixed-rate single-family mortgages purchased by Fannie Mae and Freddie Mac.

**Panel A: 30–59-day Delinquency Rate**



**Panel B: 60–89-day Delinquency Rate**



**Table I: Summary Statistics**

Panel A shows the statistics for the full sample time period (2000–2008). Panel B shows the statistics for the pre-treatment period (2000–2004) and the p-values for the t-test for difference in the control and treatment group. *Scaled applications*,  $N$ , is the number of mortgage applications in a census tract scaled by the population aged 18 to 64 years in the tract (scaled applications). *Approval ratio* (*Aprv.*) is the ratio of the number of successful applications (action type “1” or “2” in the HMDA dataset) to the number of total applications in a census tract. *Deny Credit Hist Ratio* and *Deny Debt-to-inc Ratio* are the ratio of applications denied due to credit history and debt-to-income ratio, respectively, to the number of total applications in a census tract. *Withdrawal Ratio* is the ratio of applications expressly withdrawn by the applicant to the number of total applications in the census tract.

*Economic Controls* consists of four variables described as follows: (i) *Num. Lenders* (*log*), the number of unique mortgage lenders in a census tract (expressed in natural log); (ii)  $\Delta$  *Inc per capita*, the annual growth rate of income per capita at the county level; (iii)  $\Delta$  *Emp.*, the annual growth rate of employment by all establishments at the county level; and (iv)  $\Delta$  *State GDP*, the annual growth rate of the state gross domestic product.

**Panel A: Full Sample (2000 – 2008)**

	Full Sample				Control Group (C)				Treatment Group (T)			
	N	Mean	SD	Med.	N	Mean	SD	Med.	N	Mean	SD	Med.
Scaled Applications (N)	86017	83.09	74.77	66.04	36002	98.18	77.76	77.44	50015	72.23	70.57	56.22
Approval Ratio (Aprv.)	81914	0.54	0.13	0.55	35386	0.57	0.12	0.58	46528	0.52	0.14	0.53
Deny Credit Hist Ratio	81914	0.06	0.04	0.05	35386	0.05	0.04	0.04	46528	0.06	0.05	0.05
Deny Debt-to-inc Ratio	81914	0.03	0.03	0.03	35386	0.03	0.02	0.03	46528	0.03	0.03	0.03
Withdrawal Ratio	81914	0.12	0.05	0.12	35386	0.12	0.04	0.11	46528	0.12	0.06	0.12
Num. Lenders (log)	82477	3.16	0.78	3.30	33974	3.36	0.60	3.42	48503	3.01	0.85	3.19
$\Delta$ Inc per capita	2259	0.04	0.06	0.04	1125	0.04	0.05	0.04	1134	0.05	0.07	0.04
$\Delta$ Emp	2262	0.01	0.09	0.01	1120	0.01	0.09	0.01	1142	0.01	0.10	0.01
$\Delta$ State GDP	73	0.05	0.03	0.04	29	0.05	0.02	0.04	44	0.05	0.03	0.04

**Panel B: Pre - Treatment Sample (2000 – 2004)**

	Full Sample				Control Group (C)				Treatment Group (T)				(C-T)
	N	Mean	SD	Med.	N	Mean	SD	Med.	N	Mean	SD	Med.	p-val
Scaled applications (N)	47923	110.16	83.62	92.99	20015	129.53	86.39	108.37	27908	96.27	78.68	82.84	0.000
Approval Ratio (Aprv.)	46584	0.55	0.14	0.56	19798	0.58	0.13	0.60	26786	0.52	0.14	0.53	0.000
Deny Credit Hist Ratio	46584	0.06	0.04	0.05	19798	0.06	0.04	0.05	26786	0.07	0.05	0.06	0.000
Deny Debt-to-inc Ratio	46584	0.03	0.02	0.03	19798	0.03	0.02	0.03	26786	0.03	0.02	0.03	0.000
Withdrawal Ratio	46584	0.12	0.05	0.11	19798	0.12	0.04	0.11	26786	0.13	0.05	0.12	0.000
Num. Lenders (log)	44383	3.36	0.73	3.48	17987	3.53	0.60	3.59	26396	3.24	0.78	3.39	0.000
$\Delta$ Inc per capita	1255	0.04	0.06	0.04	625	0.04	0.05	0.04	630	0.04	0.07	0.04	0.620
$\Delta$ Emp	1254	0.01	0.09	0.01	622	0.01	0.09	0.01	632	0.00	0.10	0.01	0.290
$\Delta$ State GDP	39	0.05	0.02	0.05	17	0.05	0.02	0.05	22	0.05	0.02	0.06	0.543

**Table II: Survey Evidence on the Credit Reports Usage and Discouraged Borrowers**

This table reports the regression results from the SCE Credit Access Survey. *Never* is 1 if a respondent has never checked his/her credit score (Q. N23). *Infrequently* is 1 if a respondent has never checked it or last checked it more than two years ago (Q. N23). *Unaware* is 1 if a respondent does not know his/her credit score (Q. N22). *Dscrgd* is 1 if a respondent said "I do not think I would get approved" in Q. N19. Note that this question (Q. N19) is a conditional question in the survey. Hence the observations in specifications (4–6) include only the responses in which (i) for Q. N17A, respondent selected *very unlikely* or *somewhat unlikely* to apply for mortgage/home-based loan, or refinance, or (ii) for Q. N17B, mentioned the probability of applying for mortgage or to refinance as less than 10%. All regressions include *Year*  $\times$  *Month* fixed effects (FE). Standard errors are clustered by survey's *Year*  $\times$  *Month*. p-values are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	Check Credit Report		Know Credit Score	Mortgage-discouraged Borrowers		
	(1)	(2)	(3)	(4)	(5)	(6)
	Never	Infrequently	Unaware	Dscrgd	Dscrgd	Dscrgd
Check Infrequently					0.03** (0.05)	
Unaware						0.05* (0.06)
Constant	0.08*** (0.00)	0.20*** (0.00)	0.12*** (0.00)	0.13*** (0.00)	0.13*** (0.00)	0.13*** (0.00)
Cluster (Year-Month)	Yes	Yes	Yes	Yes	Yes	Yes
FE (Year-Month)	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup>	0.007	0.007	0.007	0.003	0.004	0.005
Observations	19231	19231	20275	9059	9058	9058

**Table III: Mortgage Applications, Approval Ratio, and House Prices**

This table reports the estimates of the treatment effect of free credit reports on the number of mortgage applications, approval ratio, and growth in house prices. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

$N$ ,  $Aprv.$ , and  $\Delta HPI$  are the number of applications per 1000 adults (scaled applications), the approval ratio in a census tract and growth in house prices at census tract level, respectively. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the  $\text{Treat} \times \text{Post}$  interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All regressions include the  $\text{Border} \times \text{Year}$  fixed effects (FE) and the *Census Tract* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)
	Aprv.	Aprv.	N	N	$\Delta HPI$	$\Delta HPI$
Treat $\times$ Post	0.01*** (2.80)	0.01*** (2.84)	13.43*** (2.95)	16.63*** (3.79)	1.83* (1.88)	2.00* (1.94)
Economic Controls	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.740	0.730	0.806	0.816	0.682	0.693
Observations	81871	76437	86010	80546	24927	23806

**Table IV: Owner-occupied and Non-owner-occupied Mortgages**

This table examines the changes in (i) owner-occupied mortgage applications, (ii) non-owner-occupied applications, (iii) non-owner-occupied mortgages as the fraction of total applications, and (iv) as fraction of successful applications. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

The dependent variable in columns (1) through (4) is the number of applications per 1000 adults in a census tract (scaled applications),  $N$ . In columns (1) and (2),  $N$  measures owner-occupied category mortgage applications only; in columns (3) and (4), non-owner-occupied only. The dependent variable in columns (5) and (6) is the non-owner-occupied mortgage as a fraction of total applications, and in columns (7) and (8), as the fraction of originated applications. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat* $\times$ *Post* interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All regressions include *Border* $\times$ *Year* fixed effects (FE) and *Census Tract* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	Owner		Non-owner		Non-owner, % of all appl.		Non-owner, % of succ. appl.	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	N	N	N	N	%	%	%	%
Treat $\times$ Post	12.92*** (2.90)	15.99*** (3.75)	0.82* (1.67)	1.00* (1.80)	0.01** (2.02)	0.01* (1.74)	0.01** (2.05)	0.01* (1.91)
Economic Controls	No	Yes	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.808	0.819	0.755	0.757	0.086	0.080	0.085	0.079
Observations	86010	80546	86010	80546	81871	76437	81785	76349

**Table V: Characterizing the Effect: Heterogeneous Effects for the Overestimating Borrowers**

This table reports the treatment effect for the number of mortgage applications and approval ratio, estimated separately for the census-tract tertiles created by sorting them independently on the rejection ratios for credit history and DTI. *C. Hist.* and *DTI* respectively represent the ratio of the number of mortgage applications rejected for credit history or DTI to the total number of mortgage applications in a census tract. The tertiles for these two ratios are calculated in the pre-event year 2004. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

*N* and *Aprv.* are the number of applications per 1000 adults (scaled applications) and the approval ratio in a census tract, respectively. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat*×*Post* interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All regressions include *Border*×*Year* fixed effects (FE) and *Census Tract* FE. Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses, and the number of observations is reported in square brackets below the t-statistics. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

		N			Aprv.		
		DTI Tertiles →			DTI Tertiles →		
C. Hist. Tertiles ↓		1	2	3	1	2	3
		(1)	(2)	(3)	(4)	(5)	(6)
Treat × Post	1	20.32***	19.73***	17.41***	0.01**	0.01**	0.02**
t-statistics		(2.65)	(3.38)	(3.13)	(2.11)	(2.54)	(2.20)
Observations		[12873]	[7533]	[4518]	[12199]	[7434]	[4298]
Treat × Post	2	10.67**	9.83**	12.27**	0.01**	0.02***	0.01*
t-statistics		(2.07)	(2.44)	(2.29)	(2.20)	(3.12)	(1.77)
Observations		[7053]	[9927]	[7865]	[6956]	[9886]	[7746]
Treat × Post	3	6.23	10.47**	5.70	0.01	0.01	0.02**
t-statistics		(1.48)	(2.29)	(1.38)	(1.15)	(1.43)	(2.18)
Observations		[4955]	[7180]	[13452]	[4701]	[7145]	[13175]

**Table VI: Increase in Mortgage-related Cognizance among Borrowers**

This table reports the estimates of the treatment effect on the fraction of mortgage applications denied for credit history and debt-to-income ratio, and the in-process withdrawal ratio. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

The outcome variables are: %C.Hist, %DTI, and %WDR. %C.Hist (%DTI) is the ratio of the number of denied applications due to credit history (debt-to-income ratio) to the total number of mortgage applications in a census tract. %WDR is the ratio of number of borrower-withdrawn applications before the lender reached a decision. *High-Denial Areas* are the census tracts where denial per capita in the pre-event year 2004 was more than the regional mean of denials across the census tracts. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat*  $\times$  *Post* interaction term captures the change in the fraction of mortgage applications denied due to a given reason in the treated census tracts relative to the control census tracts. All regressions include *Border*  $\times$  *Year* fixed effects (FE) and *Census Tract* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	All Areas		High Denial Areas		All Areas		High Denial Areas		All Areas	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	% C.Hist	% C.Hist	% C.Hist	% C.Hist	% DTI	% DTI	% DTI	% DTI	%WDR	%WDR
Treat $\times$ Post	-0.003 (-1.47)	-0.003 (-1.52)	-0.003** (-2.01)	-0.003* (-1.80)	-0.002 (-1.03)	-0.002 (-1.17)	-0.002 (-1.43)	-0.002 (-1.35)	-0.009*** (-2.92)	-0.010*** (-3.95)
Economic Controls	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.541	0.532	0.575	0.568	0.267	0.264	0.319	0.320	0.340	0.338
Observations	81871	76437	38737	36447	81871	76437	38737	36447	81871	76437



**Table VII: Effect Heterogeneity by Consumer Creditworthiness**

This table reports the estimates of the treatment effect of free credit reports on the number of mortgage applications per 1000 adults, (scaled applications,  $N$ ) and the approval ratio ( $Aprv.$ ) in *ex-ante* low- and high-creditworthiness areas. A county is “subprime” if its subprime population fraction is more than the *regional mean* subprime population fraction in 1999. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

*Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the  $\text{Treat} \times \text{Post}$  interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All regressions include  $\text{Border} \times \text{Year}$  fixed effects (FE) and *Census Tract* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	<i>Ex-ante</i> High Creditworthiness (Prime Counties)				<i>Ex-ante</i> Low Creditworthiness (Subprime Counties)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	N	N	Aprv.	Aprv.	N	N	Aprv.	Aprv.
Treat $\times$ Post	16.82** (2.33)	18.80*** (2.66)	0.02*** (3.19)	0.02*** (3.29)	8.59 (1.64)	11.66** (2.42)	0.01* (1.71)	0.01* (1.78)
Economic Controls	No	Yes	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.802	0.822	0.777	0.772	0.825	0.826	0.679	0.672
Observations	39076	35703	38000	34644	46631	44558	43763	41703

**Table VIII: Effect Heterogeneity by Income Level of Consumers**

This table reports estimates of the treatment effect of free credit reports on the number of mortgage applications (Panel A) and the approval ratio (Panel B) for each of the income quartiles. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

$N$  and  $Aprv.$  are the number of applications per 1000 adults (scaled applications) and the approval ratio in a census tract, respectively. Income quartiles are calculated every year for a given census tract. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the  $\text{Treat} \times \text{Post}$  interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All regressions include  $\text{Border} \times \text{Year}$  fixed effects (FE) and *Census Tract* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

**Panel A: Number of Applications per 1000 adults**

	Income Quartile 1		Income quartile 2		Income Quartile 3		Income quartile 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	N	N	N	N	N	N	N	N
Treat $\times$ Post	0.23 (0.16)	0.49 (0.35)	2.14** (2.54)	2.38*** (3.11)	2.64** (2.35)	3.03*** (3.19)	3.97* (1.90)	4.83*** (2.61)
Economic Controls	No	Yes	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.760	0.760	0.772	0.773	0.740	0.741	0.659	0.672
Observations	87479	80546	87479	80546	87479	80546	87479	80546

**Panel B: Approval Ratio**

	Income Quartile 1		Income quartile 2		Income Quartile 3		Income quartile 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Aprv.	Aprv.	Aprv.	Aprv.	Aprv.	Aprv.	Aprv.	Aprv.
Treat $\times$ Post	0.01* (1.92)	0.01** (2.24)	0.01 (1.24)	0.01 (1.18)	0.00 (0.37)	0.00 (0.34)	-0.00 (-0.41)	0.00 (0.19)
Economic Controls	No	Yes	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.316	0.307	0.338	0.326	0.308	0.297	0.169	0.156
Observations	71190	66014	71718	66523	71832	66632	71248	66062

**Table IX: First-time Homebuyers and Mortgage Interest Rates in the GSE Data**

This table reports the estimates of the treatment effect on the fraction of first-time homebuyers and interest rate using the GSE data. In column (1) and (2), the dependent variable is the ratio of the number of mortgages taken by first-time homebuyers to total number of mortgages for which the information on first-time homebuyers is not missing, calculated at the zip3-state area level. The regression specification is from Equation (3):

$$Y_{zsjt} = \beta_0 + \beta_1 \text{Treat}_{zsj} \times \text{Post}_t + \delta \times \text{Economic controls} + \alpha_{zs} + \gamma_{jt} + \varepsilon_{zsjt}.$$

In columns (3) and (4), the dependent variable is interest rate on the GSE mortgages (in percentages), and the regression specification is from Equation (4):

$$\text{Interest Rate}_{izsjt} = \beta_0 + \beta_1 \text{Treat}_{izsj} \times \text{Post}_t + \delta \times \text{Controls} + \alpha_{zs} + \gamma_{jt} + \varepsilon_{izsjt}.$$

*Economic Controls* include annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). *Mortgage Controls* refer to *credit score* and *combined loan-to-value (CLTV)* in column (3), and to three additional controls, *DTI ratio*, *number of units in the property*, and *mortgage insurance percentage*, in column (4). The coefficient associated with the *Treat*×*Post* interaction term captures the change in the dependent variable in the treated zip3-state areas vis-a-vis the control. All regressions include *Zip3–State* fixed effects (FE) and *Border*×*Quarter* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	First-time Borrower Ratio		Interest Rate (%)	
	(1)	(2)	(3)	(4)
Treat × Post	0.011*** (2.70)	0.010** (2.47)	0.009*** (13.36)	0.010*** (12.06)
Economic Controls	No	Yes	-	-
Mortgage Controls	-	-	Yes	Yes
Zip3-State FE	Yes	Yes	Yes	Yes
Border × Qtr FE	Yes	Yes	Yes	Yes
Cluster Zip3-State	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.694	0.695	0.731	0.758
Observations	7593	7593	7579052	3512619
Reg. Unit	Zip3-state Aggregate	Zip3-state Aggregate	Individual Mortgage	Individual Mortgage

**Table X: Effect Heterogeneity by Lenders Density**

This table reports the estimates of the treatment effect on the origination volume (in 1000 USD) per adult and the approval ratio, estimated separately for census tracts having a high and low density of mortgage lenders per capita in 2004. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

*Low (High)* identifies a census tract having a lower (higher) number of HMDA lenders than the *regional mean* number of HMDA lenders (per census tract) within the bordering counties between the given control state and all the treatment states surrounding it in 2004 (See Footnote 21). *Difference [High - Low]* shows the result of the t-test for the difference in coefficients of *Treat* × *Post* in specifications *High* and *Low*. The dependent variable in Columns (1) through (4) is volume of mortgages originated (in 1000 USD) per adult in a census tract. The dependent variable in Columns (5) through (8) is the approval ratio of mortgage applications at census tract level. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat* × *Post* interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	Volume (in 1000 USD) per Adult				Approval Ratio			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Low	High	Low	High	Low	High	Low	High
Treat × Post	0.002** (2.21)	0.001 (1.16)	0.003*** (3.06)	0.002* (1.66)	0.015*** (3.07)	0.010* (1.92)	0.016*** (3.15)	0.009* (1.73)
Difference [High - Low]		-0.001		-0.001		-0.006		-0.007
p-value		(0.592)		(0.498)		(0.474)		(0.413)
Economic Controls	No	No	Yes	Yes	No	No	Yes	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Border × Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.643	0.571	0.636	0.616	0.751	0.716	0.740	0.709
Observations	60210	25497	56188	24073	57134	24629	53135	23212

**Table XI: Did Origination Increase due to Rise in Private Mortgage Securitization?**

This table reports the estimates of the treatment effect on the approval ratio estimated separately for mortgages sold to non-GSEs, sold to GSEs, and not sold. The regression specification is from Equation (1):

$$Y_{icsjt} = \beta_0 + \beta_1 \text{Treat}_{icsj} \times \text{Post}_t + \delta \times \text{Economic Controls} + \alpha_i + \gamma_{jt} + \varepsilon_{icsjt}.$$

The dependent variables are the fraction of total mortgage applications originated and sold to the non-GSEs (Columns 1 and 2); originated and sold to the GSEs (Columns 3 and 4); approved and not sold by the lending institution (Columns 5 and 6). All the dependent variables are calculated at the census tract level. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat*  $\times$  *Post* interaction term captures the change in the dependent variable in the treated census tracts relative to the control census tracts. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	Sold to Non-GSE		Sold to GSE		Not Sold	
	(1)	(2)	(3)	(4)	(5)	(6)
	Fraction	Fraction	Fraction	Fraction	Fraction	Fraction
Treat $\times$ Post	-0.004 (-0.28)	0.001 (0.05)	0.048** (2.49)	0.047*** (2.78)	0.001 (0.11)	0.002 (0.52)
Economic Controls	No	Yes	No	Yes	No	Yes
Census Tract FE	Yes	Yes	Yes	Yes	Yes	Yes
Border $\times$ Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (County)	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.008	-0.003	0.003	-0.003	0.055	0.028
Observations	81871	76437	81871	76437	81871	76437

**Table XII: Did Origination Increase due to Subprime Lending? Credit Score-based Evidence**

This table reports the estimates of the treatment effect on the number of mortgages originated to prime and subprime borrowers by government sponsored enterprises (GSEs) Fannie Mae and Freddie Mac. The regression specification is from Equation (3):

$$Y_{zsjt} = \beta_0 + \beta_1 \text{Treat}_{zsj} \times \text{Post}_t + \delta \times \text{Economic controls} + \alpha_{zs} + \gamma_{jt} + \varepsilon_{zsjt}.$$

The dependent variable in Column (1) is *N-Prime*, the number of mortgages originated to prime borrowers (credit score  $\geq 620$ ) in a given zip3-state area. The dependent variable in Column (2) is *N-Subprime*, the number of applications to subprime borrowers (credit score  $< 620$ ) in a given zip3-state area. *Economic Controls* include the number of mortgage lenders in a census tract and annual growth rate of county income per capita, county aggregate employment, and state gross domestic product (GDP). The coefficient associated with the *Treat*  $\times$  *Post* interaction term captures the change in the dependent variable in the treated zip3-state areas relative to the control zip3-state areas. All regressions include *Zip3-State* fixed effects (FE) and *Border*  $\times$  *Quarter* FE. All variables are defined in Table (I). Standard errors are clustered by county. t-statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

	(1)	(2)	(3)	(4)
	N-Prime	N-Prime	N-Subprime	N-Subprime
Treat $\times$ Post	325.63*** (3.54)	325.87*** (3.44)	11.28** (2.24)	11.40** (2.23)
Economic Controls	No	Yes	No	Yes
Zip3-State FE	Yes	Yes	Yes	Yes
Border $\times$ Qtr FE	Yes	Yes	Yes	Yes
Cluster Zip3-State	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.761	0.762	0.795	0.796
Observations	7599	7599	7599	7599

**Table XIII: Effect of Free Credit Reports on Banks**

Panel (A) of this table reports the estimates of the treatment effect on financial performance of banks. Panel (B) reports the estimates of the originated mortgage amounts (in million \$) separately estimated for the banks with an *ex-ante* low and high share of liquid assets. The regression specification is from Equation (5):

$$Y_{bt} = \beta_0 + \beta_1 \text{Treat}_b \times \text{Post}_t + \delta \times \text{Bank Controls}_{bt} + \alpha_l + \gamma_t + \varepsilon_{bt}.$$

$t$  refers to *Year–Quarter* in Panel (A) and *Year* in Panel (B). *NIM* (Net Interest Margin) is the ratio of net interest income to earning assets (in percentages); *RoE* (Return on Equity) is the ratio of net income to book value of equity (in percentages); and *RoA* (Return on Assets) is the ratio of net income to book value of total assets (in percentages). A bank is classified as having an *Ex-ante High Liq. Share* if its share of liquid assets (of their total assets) in the pre-event year 2004 was *greater* than the cross-sectional sample mean, and as having an *Ex-ante Low Liq. Share* otherwise. Bank Controls in Panel (A) include: natural log of the total assets (in \$1000); cost of deposit (ratio of total interest expense to total earning assets, expressed in percentages); and share of liquid assets in total assets (in percentages). Bank Controls in Panel (B) include only the first two of these variables. The coefficient associated with the *Treat* × *Post* interaction term captures the change in the dependent variable for the treated banks relative to the control banks. All regressions in Panel (A) include *Year–Quarter* fixed effects (FE) and *Bank* FE; and those in Panel (B) include *Year* FE and *Bank* FE. Standard errors are clustered by county.  $t$ -statistics are reported below the coefficients in parentheses. \*, \*\*, and \*\*\* denote statistical significance at the 10%, 5%, and 1% level, respectively.

**Panel A: Financial Performance of Banks**

	(1)	(2)	(3)	(4)	(5)	(6)
	NIM (%)	NIM (%)	RoE (%)	RoE (%)	RoA (%)	RoA (%)
Treat × Post	0.06*** (5.49)	0.06*** (5.96)	0.75*** (5.13)	0.76*** (5.32)	0.07*** (5.18)	0.08*** (5.53)
Bank Controls	No	Yes	No	Yes	No	Yes
Bank FE	Yes	Yes	Yes	Yes	Yes	Yes
Year-Qtr FE	Yes	Yes	Yes	Yes	Yes	Yes
Cluster (Bank)	Yes	Yes	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.807	0.814	0.586	0.597	0.556	0.573
Observations	85847	85847	85847	85847	85847	85847

**Panel B: Mortgage Origination by *Ex-ante* Share of Liquid Assets of Banks**

	<i>Ex-ante</i> Low Liq. Share		<i>Ex-ante</i> High Liq. Share	
	(1)	(2)	(3)	(4)
	Amt	Amt	Amt	Amt
Treat × Post	-159.18 (-1.29)	-157.24 (-1.28)	111.33* (1.67)	111.35* (1.69)
Bank Controls	No	Yes	No	Yes
Bank FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Cluster (Bank)	Yes	Yes	Yes	Yes
R <sup>2</sup> (Adj.)	0.569	0.570	0.792	0.792
Observations	10940	10940	7748	7748



# Appendix

## Data Appendix

The HMDA data contain 190.4 million mortgage applications over the sample period (2000–2008). These application-level data were aggregated to the *Census Tract* × *Year* panel in several steps. First, all observations that had state, county or census tract information missing or “NA”, or state Federal Information Processing Standard (FIPS) code of “0”, “00” or “0 ” were dropped (2.5% of the observations), leaving 185.6 million mortgages with an identifiable county. Then, observations on three action types were removed: covered loans purchased by the financial institutions from other institutions (18.80%), as these are not borrower initiated; pre-approval requests denied by financial institutions (0.01%), as these data were included in HMDA reporting only from 2004; and pre-approval requests approved by the financial institutions but not accepted by the applicants, as these data, too, were included in the HMDA reporting only from 2004, and this reporting is not mandatory (0.025%). This leaves 150.7 million applications belonging to 77,526 unique census tracts (603,849 *Census Tract* × *Year* observations). Finally, with the help of the county adjacency data from the [Census Bureau \(n.d.\)](#), those census tracts that belong to the bordering counties of the treated and control states were selected. This led to the HMDA regression sample: 89,535 *Census Tract* × *Year* observations consisting of 11,942 unique census tracts of which 7,011 are treated and 4,931 are control.

The GSE data contain 33 million observations over the sample period. The property locations in this data do not contain the census tract information, but only the first 3 digits of the zip code (zip3) and state. Hence, to identify the mortgages from the zip3-states that lie within the bordering counties of the sample, the zip code-to-county crosswalk file provided by the U.S. Department of Housing was used.<sup>28</sup> Then, aggregating the individual mortgages to the zip3-state level and restricting the sample to only those zip3-states that lie within the sample border counties yielded 221 unique zip3-states (91 control and 130 treated) and 7,599 *Zip3-State* × *Quarter* observations.

Finally, the mortgage lenders in the HMDA data were matched with the commercial banks in the Call Reports (FFIEC Forms 031/041) data using lenders’ Federal Deposit Insurance Cor-

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<sup>28</sup> Areas delimited by 3-digit zip codes do not align with the county borders. Hence, to identify the 3-digit zip codes that lie along the county borders, first, a crosswalk file of 5-digit zip codes to county is obtained from the [Office of Policy Development and Research \(n.d., 2010 Q1 version\)](#). Then all such 3-digit zip codes are filtered out from the sample for which none of the underlying 5-digit zip codes lie within the bordering counties.

poration (FDIC) certificate ID, or Office of the Comptroller of the Currency (OCC) charter number (henceforth, the identifiers). Call Reports contain information on banks' identifiers and also a unique id called RSSD ID. At the same time, HMDA data contain a lender's agency code (lender's regulator) and a respondent ID. A respondent ID equals the FDIC Certificate ID if the lender's regulator is the FDIC; and it equals the OCC charter number if the regulator is the OCC.

Some HMDA mortgage lenders are the affiliates of the commercial banks, but are not banks themselves. Such lenders were matched using their parent entities (available in the HMDA Ultimate Panel data). If both an HMDA reporter and its parent entity had a successful match in the call reports, the parent's match was kept. Finally, the RSSD ID began to be directly available in the HMDA data from 2004, so the matching was done for subsequent years using this ID, instead of the combination of the agency code and respondent ID.

# **KNOW THYSELF: FREE CREDIT REPORTS AND THE RETAIL MORTGAGE MARKET**

**Online Appendix**

**Table A1: Sample States and Status of Selected Regulations**

This table lists all the states included in the sample and their treatment and control status. For each state, it also indicates whether mortgages are recourse or non-recourse ([Ghent & Kudlyak, 2011](#)), whether foreclosures are judicial or non-judicial ([Gerardi et al., 2013](#)), and whether and when the state adopted Anti-predatory Lending laws (APL) ([Di Maggio & Kermani, 2017](#)).

State	Treatment (T) / Control (C)	Recourse (R) / Non-Recourse (NR)	Judicial (J) / Non-judicial (NJ)	APL (Adoption Month, Year) / Non-APL (NAPL)
Alabama	T	R	NJ	NAPL
Arizona	T	NR	NJ	NAPL
Colorado	C	R	NJ	APL (Jul, 2003)
Connecticut	T	R	J	APL (Jan, 2002)
Delaware	T	R	J	NAPL
Florida	T	R	NJ	NAPL
Georgia	C	R	J	APL (Mar, 2003)
Kansas	T	R	J	NAPL
Maryland	C	R	J	APL (Oct, 2002)
Massachusetts	C	R	J	APL (Nov, 2004)
Nebraska	T	R	J	NAPL
New Hampshire	T	R	NJ	NAPL
New Jersey	C	R	J	APL (Nov, 2003)
New Mexico	T	R	J	APL (Jan, 2004)
New York	T	R	J	APL (Apr, 2003)
North Carolina	T	NR	NJ	APL (Jul, 2000)
Oklahoma	T	R	NJ	NAPL
Pennsylvania	T	R	J	NAPL
Rhode Island	T	R	NJ	APL (Dec, 2006)
South Carolina	T	R	J	APL (Jan, 2004)
Tennessee	T	R	NJ	NAPL
Utah	T	R	NJ	NAPL
Vermont	C	R	J	NAPL
Virginia	T	R	NJ	NAPL
West Virginia	T	R	J	APL (Jun, 2000)
Wyoming	T	R	NJ	NAPL