Credit Standards and Segregation[*]

Amine Ouazad[†] Romain Rancière[‡]

January 2014

Abstract

This paper explores the effects of changes in lending standards on racial segregation within metropolitan areas. Such changes affect neighborhood choices as well as aggregate prices and quantities in the housing market. Using the credit boom of 2000-2006 as a large-scale experiment, we put forward an IV strategy that predicts the relaxation of credit standards as the result of a credit supply shock predominantly affecting liquidity-constrained banks. The relaxed lending standards led to significant outflows of Whites from black and from racially mixed neighborhoods: without such credit supply shock, black households would have had between 2.3 and 5.1 percentage points more white neighbors in 2010.

[*] We thank Patrick Bayer, Leah Platt Boustan, Jess Benhabib, Christopher Crow, Giovanni Dell'Ariccia, Susan Dynarski, Denis Gromb, Maria Guadalupe, Brian Jacob, Robert Hunt, Jean Imbs, Matthew Kahn, Luc Laeven, Ross Levine, Alan Manning, Loretta Mester, Atif R. Mian, Guy Michaels, Max Nathan, Joel Peress, Gordon Philipps, Thomas Piketty, Steve Pischke, Rodney Ramcharan, Stephen Ross, Emmanuel Saez, Jose Scheinkman, John Van Reenen, Dubravka Ritter, Aaron Tornell, and Ekaterina Zhuravskaya for discussions and helpful suggestions on preliminary versions of this paper as well as Elena Loutskina and Luc Laeven for sharing some data. For insightful comments we thank the audiences at University of Amsterdam, Duke, Georgetown, LSE, UCLA, Michigan, PSE, INSEAD, RAND, Syracuse, the Federal Reserve Bank of Philadelphia, the CEPR Public Policy Symposium, Society of Labor Economists conference, the Urban Economics Conference, the Econometric Society, and the Rimini conference meeting. The usual disclaimers apply. We thank INSEAD, the International Monetary Fund, and the Paris School of Economics for financial and computing assistance.

[†] INSEAD.

[‡] International Monetary Fund and Paris School of Economics.

1. INTRODUCTION

The availability and affordability of mortgage credit is a key determinant of housing choices. Large aggregate changes in mortgage lending standards, as was experienced during the last mortgage credit boom, could thus have large effects on the sorting of households by income, race, or education across neighborhoods. This paper focuses on the role played by credit market conditions on the dynamics of urban segregation, using the last U.S. mortgage credit boom as a large-scale experiment.

While there is a vast literature on the determinants of urban racial segregation (e.g., Bayer, McMillan, and Rueben 2004; Cutler, Glaeser, and Vigdor 2008), the role played by mortgage credit standards in shaping aggregate racial segregation has so far received little attention.[1] Some have suggested that the last mortgage credit boom could have contributed to the decline in segregation over the last decade.[2] Yet, the effect of a change in lending standards on metropolitan area segregation has, to the best of our knowledge, never been formally tested.

Minorities are generally considered to be more credit constrained than other groups (Ross and Yinger, 2002) and thus minorities could be expected to benefit more from relaxed lending standards. With an increased availability of mortgage credit, minority households have access to a larger set of housing options. These include the possibility to relocate into more racially mixed neighborhoods but also into neighborhoods with a comparable racial mix but with more desirable characteristics. This partial equilibrium perspective ignores, however, the role of general equilibrium effects. An increase in the supply of mortgage credit affects how the preferences for neighborhoods and the preferences for rental vs. homeownership *of all households* translate into actual housing decisions. In general equilibrium, these decisions lead to changes in housing prices, neighborhood demographics, and the supply of housing. Changes in credit market conditions can therefore lead to either a decline or an increase in urban segregation.

---

[1] This paper's focus on metropolitan level changes in segregation due to market forces is key – there is an extant literature on discrimination in mortgage lending and redlining (Ross and Yinger 2002).

[2] *"Several of the metropolitan areas with the greatest declines in segregation are also areas associated with significant exposure to the subprime mortgage market. It is also true that several metro areas with significant subprime exposure—such as Miami and Las Vegas—appear to have followed fairly unremarkable segregation trajectories over the past decade."* (Glaeser and Vigdor 2012)

The objective of this paper is twofold. First, we design an empirical strategy to identify the causal effect of the relaxation of mortgage lending standards on racial segregation across neighborhoods at the metropolitan area level during the recent credit boom. Second, we use additional micro-level data to show how credit supply affects population flows by race, and how these flows are facilitated or hindered by general equilibrium changes in the relative price of housing in neighborhoods with different racial compositions within a metropolitan area.

We examine the impact of credit standards on segregation by combining information from the universe of mortgage loan applications[3], made publicly available through the Home Mortgage Disclosure Act, with Census-based information on racial demographics. The mortgage credit boom of 2000 to 2006 saw large changes in mortgage applications' approval rates by lenders and in mortgage borrowers' loan-to-income (LTI) ratios. While banks approved 70% of mortgage applications in 2000, this rate jumped to 84% in 2006, a 14 percentage point jump that is comparable to the increase for Hispanics (13 percentage points) and Blacks (13.1 percentage points). In 2000, the average new homeowner borrowed 1.9 times his income, whereas by 2004 this ratio had risen to 2.4 times his annual income.

Using demographic information at the census tract level, we construct a set of standard measures of metropolitan-area level racial segregation across tracts, in 2000 and 2010. OLS estimation results show that in metropolitan areas that experienced larger increases in LTI ratios and mortgage loan approval rates during the credit boom of 2000 to 2006, black households had fewer white neighbors – a decline in black exposure to Whites – and black segregation declined slower than in other metropolitan areas. Although this positive correlation between lending standards relaxation and black segregation is intriguing, there are significant challenges when identifying the *causal* impact of relaxing mortgage lending standards on segregation. Observed approval rates and loan-to-income ratios – two measures of lending standards – reflect both supply and demand factors.

This paper's identification strategy relies on instruments that identify the relaxation of lending standards due to an increase in credit supply separately from an increase in credit demand. The instrumental variables are measures of banks' liquidity conditions at the metropolitan level in the

[3] We use the sample of mortgage applications for single-family owner-occupying purchases, excluding loans by the Federal Housing Administration.

early 1990s, that is, prior to a number of key transformations that affected the mortgage industry and favored the rise of securitization. The underlying hypothesis is that increased securitization activity allowed banks with initially low levels of liquidity to *catch up* by increasing their approval rate and median loan-to-income ratios relative to banks with initially high levels of liquidity. First stage estimation results show that indeed metropolitan areas with a low level of bank liquidity in 1990-1994 exhibited both a greater relaxation of lending standards, and a higher growth of mortgage securitization volumes, between 2000 and 2006, than metropolitan areas with an initially high level of liquidity. Bank liquidity in the early 1990s is a strong predictor of future changes in approval rates and in originations' loan-to-income ratio during the boom, yet it is uncorrelated with observable factors (most importantly, Hispanic inflows, and income changes on average and by racial group) affecting mortgage credit demand. In addition, we find that our instruments do not display any significant correlation with a number of other potential confounders affecting racial segregation such as local amenities, the level of crime, the college premium, and income inequality.

Instrumental variable (IV) regression results suggest that the decline in lending standards during the boom had a robust and significant effect on segregation. The effect is economically important: in our estimations, the magnitude of the boom's observed increase in loan-to-income ratios (resp. approval rates) lowered the fraction of white neighbors in the tract of an average black resident by 5.1 (resp. 2.3) percentage points, while having no significant and robust impact on the fraction of Hispanic neighbors in the tract of an average black household. Given the decline of segregation during the last decade[4], our results suggest that the increased supply of credit *slowed down* the racial integration of cities**.**

The paper's findings survive a series of robustness tests. In particular, census-based measures of segregation are decennial; hence post-boom (2007–2010) credit conditions could substantially alter our conclusions. However, controlling for metropolitan area 2007-2010 foreclosure rates leaves our main estimates statistically and economically unchanged. Additionally, two data sets provide racial segregation measures for 2006 and 2008: (i) the 2006-2010 American Community Survey (ACS) provides tract-level data for households interviewed between 2006 and 2010, in which the median household responded in 2008, two years before the

[4] See Glaeser and Vigdor (2012) .

2010 Census; results based on these data are very similar to those based on Censuses. (ii) The Department of Education's *annual* school demographics allow us to compute metropolitan area measures of racial segregation across schools in 2000 and 2006. The effect of credit standards relaxation on *school* segregation is strongly similar to its effect on *urban* segregation.

Our metropolitan area findings could result from black mobility into black neighborhoods or from white mobility out of black neighborhoods. We show using census tract level data that the decline in lending standards has contributed to foster significant white mobility out of both racially mixed and mostly black census tracts, i.e. tracts with between 10% and 60% black population, and tracts with at least 60% black population, and into mostly white census tracts (tracts with less than 10% of black population). This mobility pattern with outflows starting at about 10–15% of black population are consistent with a tipping point model (Card, Mas, and Rothstein 2008). In addition, we find that lending standards relaxation led to significant black residents mobility into racially mixed tracts, which were experiencing whites outflows, but did not lead to black mobility into tracts that were mostly white in 2000.

General equilibrium effects may explain such lack of black mobility towards mostly white tracts in metropolitan areas that experienced a decline in lending standards. Indeed, the simple model of neighborhood choice with endogenous prices and borrowing constraints, presented in the appendix, suggests that lending standards relaxation leads to an increase in prices in desirable neighborhoods, potentially pricing out minorities from such neighborhoods. In our model, an increase in credit supply leads to an increase in segregation whenever the partial equilibrium effect (the effect at given prices) is offset by the general equilibrium effect of prices on segregation. Empirically, we find that, during the boom, house prices increased significantly more in mostly white tracts than in either racially mixed or mostly black tracts, further hindering black households' mobility into such tracts, but only in metropolitan areas that experienced a significant decline in lending standards. This price result within metropolitan areas, across tracts, holds even after controlling for migrations, foreclosures, and metropolitan area house price changes.

Indeed, such micro-level evidence on household mobility and price increases suggests that metropolitan areas with less elastic housing supply might have experienced a stronger effect of lending standards on racial segregation. At the metropolitan area level, using the housing supply elasticity measures of Saiz (2008), we indeed find that the positive impact of the relaxation of lending standards on segregation is much stronger in metropolitan areas with low housing supply elasticity.

The rest of the paper proceeds as follows. Section 2 discusses related literature. Section 3 presents the main data sources, the evolution of segregation in the past decade, and the change in credit conditions during the boom; it also describes the observed strong correlations between changes in segregation and changes in credit conditions. Section 4 presents the instrumental variables strategy and the paper's main results, and discusses its robustness. Section 5 identifies economic mechanisms that are consistent with a simple model of neighborhood choice with credit constraints. Section 6 concludes.

2. RELATED LITERATURE

In 2000, a majority of urban Blacks lived in highly segregated neighborhoods (Massey 2004), and ample evidence suggests that racial segregation has negative impacts on black welfare (Cutler and Glaeser 1997) in a number of dimensions: education (Card and Rothstein 2007), black well-being (Massey, Condran, and Denton 1987), and labor market opportunities of black youth (Raphael 1998). Analysis of racial segregation in the first half of the 20th century stressed the importance of white households' collective action, including land use regulations and racial covenants, in shaping the distribution of races across neighborhoods (Cutler, Glaeser, and Vigdor 1999; Brooks 2011). Literature on racial segregation in the second half of the 20th century, both theoretical and empirical, model segregation across neighborhoods as the result of an equilibrium in which large price differences across white and minority neighborhoods reflect differences in the quality of local amenities (Epple and Sieg 1999), differences in households' income and education (Benabou 1996), and white households' preference for same-race neighbors (Krysan and Farley 2002). The mobility of white households away from black neighborhoods, also called white flight, described in Schelling's (1971) seminal work, was facilitated by declining transportation costs (Baum-Snow 2007), triggered by the presence of minorities, and by inequalities in the quality of public services and education across neighborhoods (Boustan 2010). This paper's main goal is to contribute to this literature by understanding whether large changes in mortgage credit markets hinders or facilitates the mobility of white and minority households across neighborhoods.

The nature and extent of racial discrimination in mortgage loan approvals has been estimated using both observational data with a large range of covariates (Munnell et al. 1996) and audit pair studies (Ross and Yinger 2002). Literature suggests that banks may have conditioned their mortgage approval decisions on the neighborhood's racial composition, a phenomenon called 'redlining' (Ross and Tootell 2004). Such lender behavior limits minorities' opportunities to relocate in more racially diverse neighborhoods. The estimated magnitude of racial discrimination in mortgage approvals can be large, but their interpretation as reflecting banks' behavior rather than minority applicants' unobservables remains controversial.

This paper stresses the possibility of large impacts of mortgage lending standards on racial segregation even absent significant racial discrimination in mortgage lending. In particular, recent analysis of the credit boom has suggested that the credit boom saw a large increase in house prices (Himmelberg, Mayer, and Sinai 2005; Campbell, Davis, Gallin, and Martin 2009) as well as a large increase in the dispersion of prices across neighborhoods (Gyourko, Mayer, and Sinai 2006). Evidence also suggests that the later period of the housing boom (2000-2006) saw substantial population flows across neighborhoods (Guerrieri, Hartley, and Hurst 2012). Because of such large changes in the relative prices and the demographic composition of neighborhoods across time, a relaxation of lending standards, which increases applicants' neighborhood choice set *at given prices*, may actually *reduce* applicants' choice set, when accounting for neighborhood price changes and demographic flows. The simple model presented in the appendix formalizes this last point.

Identifying the impact of changes in lending standards, as opposed to changes in the demand for credit, is subtle, in large part because of the simultaneity problem: median loan-to-income ratios and approval rates reflect changes in both the demand and the supply of credit. This paper is focused on identifying the impact of the latter on metro-level racial segregation. The identification strategy we adopt here predicts changes in mortgage lending standards during the boom at the metropolitan area level using banks' balance sheet structure in the early 1990s. Focusing on the supply side of the credit market is well in line with recent literature showing that credit supply – e.g., through more lenient lending standards – is responsible for a large share of the rise in leverage and approval rates during the credit boom. Mian and Sufi (2009) use ZIP code level data to demonstrate that a supply-based channel is the most likely explanation for the mortgage market's expansion during the boom. Dell'Arriccia, Igan, and Laeven (2009) document using Home Mortgage Disclosure Act application data that growing origination volumes were correlated with relaxed lending standards. Favara and Imbs' (2010) paper is another piece of evidence that confirms the role of credit supply, as they show how the timing and extent of interstate banking and branching deregulation increased loan volume and LTI ratios, and led to falling denial rates.[5] Keys, Mukherjee, Seru, and Vig (2010) demonstrate how securitization led to both increasing mortgage credit supply and declining lending standards. This securitization boom led to the development of the originate-to-distribute model, which, as Purnanandam (2011) shows, strongly benefited capital-constrained banks. However, we do not exclude that changes in housing or credit demand partly drove the rise in credit supply and the relaxation of credit standards. In particular, Ferreira and Gyourko (2011) use disaggregated census tract level data and argue that local income shocks preceded increases in property prices. Hence, an important identification challenge in the literature is to identify the impact of lenders' underwriting policies separately from the impact of income changes, when estimating the impact of credit supply on racial segregation.

Our new identification strategy provides instruments that predict loan-to-income ratio and approval rate changes at the metropolitan area level. We will show how these instruments can be used both in metro-level IV regressions, to identify the causal impact of a relaxation of lending standards on segregation (Section 4), and at the census-tract level to analyze how demographic changes between neighborhoods within metro areas vary with the predicted changes in lending standards (Section 5).

3. Data Set and Descriptive Evidence

3.1 Data Sources

We use mortgage data for the years 1995–2007 that was compiled in accordance with the Home Mortgage Disclosure Act (HMDA), which mandates reporting by most depository and non-depository lending institutions.[6] HMDA disclosure requirements thus apply to more than 90% of all mortgage applications and originations (Dell'Arriccia, Igan, and Laeven 2009), and for each mortgage lender report the loan amount, the applicant's income, the applicant's race and gender,

[5] Ng (2012) and Adelino, Schoar, and Severino (2012) provide additional empirical support for a supply-driven credit boom.
[6] Specifically, HUD regulates for-profit lenders that have combined assets exceeding $10 million and/or originated 100 or more home purchase loans (including refinancing loans) in the preceding calendar year.

and the census tract of the house. We focus on credit standards for single-family, owner-occupied mortgages.

The Census Bureau's Summary File I provides census tract–level demographics for the 2000 and 2010 censuses. We construct measures of racial demographics and racial segregation across census tracts for each metropolitan area, following measures described in Cutler, Glaeser, and Vigdor (1999) and Massey, White, and Phua (1996). We equate "metropolitan areas" with the widely used Core Based Statistical Areas (CBSAs) in 2003 borders.[7] CBSAs encompass both metropolitan statistical areas (MSAs) and "micropolitan" statistical areas (μSAs). [8]

The banks' balance sheet data used to compute our liquidity measures come from the Federal Reserve's Reports of Condition and Income, also known as Call Reports. As explained in detail in section 4, these balance sheet data will be merged with HMDA data on mortgage origination by banks in order to produce our MSA-level measures of liquidity (our instrument) in the early 1990s.

3.2 Racial Segregation from 2000 to 2010

From the many available segregation measures (Massey and Denton 1988) we choose the isolation and exposure indices, which have been extensively used in the literature (Cutler, Glaeser and Vigdor 1999). Here *isolation* is defined as the average fraction of neighbors of the same race in the average census tract of Whites, Blacks, or Hispanics;[9] thus, the isolation of Whites is the average fraction of white neighbors for white households. The isolation index is an especially relevant measure when considering the effect of neighbors on outcomes – as in standard models with a linear-in-means peer effects specification (Manski 1993), where average peers' characteristics are the main contextual input for considering either neighborhood-level (Goux and Maurin 2007) or school-level social interactions (Hoxby 2001). We focus on isolation

[7] We keep consistent metropolitan area borders throughout the dataset, following the 2003 definitions.

[8] For clarity and simplicity we refer to 'metropolitan areas.' The Census Bureau defines two kinds of metropolitan areas, metropolitan statistical areas (MSAs) and micropolitan statistical areas (μSAs). A *metropolitan statistical area* is a contiguous geographic area containing a large population core (of more than 50,000 inhabitants) and adjacent communities that are highly integrated (as measured by commuting time) with that core. The concept of a *micropolitan statistical area* parallels that of the MSA but with a lower core threshold (i.e., more than 10,000 inhabitants). Our metropolitan areas include both MSAs and μSAs.

[9] To keep the discussion manageable, we do not display results for Asians, American Indians and Alaska Natives, or Pacific Islanders. However, our findings for Asians suggest significant effects for this group (results available upon request).

for clarity, but our results are robust to using instead either dissimilarity[10] or normalized isolation (Cutler, Glaeser, and Vigdor 1999).[11]

The isolation of Whites in metropolitan area k is expressed formally as

$$Isolation_k(whites) = \sum_j \frac{whites_{k,j}}{whites_k} \frac{whites_{k,j}}{population_{k,j}}$$

where $whites_{k,j}$ is the white population in census tract j of metropolitan area k; $whites_k$ is the overall white population in metropolitan area k; and $population_{k,j}$ is the total population in census tract j of metropolitan area k.

White isolation decreases as white households are more exposed to neighbors of other races. For instance, the exposure of Whites to Blacks in metropolitan area k may be written as

$$Exposure_k(whites|blacks) = \sum_j \frac{whites_{k,j}}{whites_k} \frac{blacks_{k,j}}{population_{k,j}}$$

where $blacks_{k,j}$ is total black population in census tract j of metropolitan area k. In the case of two racial groups, one group's isolation increases as exposure to other group decreases.

From 2000 to 2010, black and white racial segregation across census tracts continued its well-documented decline that began in the 1970s (Glaeser and Vigdor 2012), as shown in Table A1 of the appendix. Black isolation declined in about three quarters of the Metropolitan Statistical Areas (MSAs). In the average metropolitan area, in 2000, the average black resident lived in a census tract for which 50.5% of the population was of the same race (i.e., black isolation was 50.5%); this same fraction declined to 45.4% in 2010. However, black exposure to whites *declined* over the period in 79 percent of the MSAs, with a median reduction of 2.0 percentage points. Thus, the decline in black isolation is largely explained by the increased exposure of black residents to Hispanics, which occurs in almost all metropolitan areas (98.1%); on average black residents live with 3.7 percentage points more Hispanic neighbors in 2010 than in 2000.

[10] So, for example, the *dissimilarity* of Blacks is the fraction of Blacks that would need to move in order to yield an even distribution of Blacks across census tracts. The *normalized isolation* of Blacks is the isolation of Blacks minus the FBIM *divided by* 1 minus the FBIM, where for notational convenience the acronym stands for "fraction of Blacks in the metro area".

[11] The normalized isolation captures some mechanical demographic changes. We use the non-normalized isolation index and control for demographic changes in the main regression, thereby retaining a more natural interpretation for the coefficients of interest.

3.3 Credit Conditions from 2000 to 2006

We measure overall credit conditions by the median loan-to-income ratio and the mortgage application approval rate (i.e., 100% minus the denial rate).[12] The median LTI ratio captures the extent to which a typical borrower can leverage her income. We explain our particular choice of credit standard measures in this section.

The median LTI for the entire population of mortgage originations increased from 1.89 to 2.3 during this period, with similar upward trends for the three major racial groups: an increase of 0.40 for white borrowers, 0.42 for black borrowers, and 0.41 for Hispanic borrowers.[13] It should be emphasized that, whereas the loan-to-*income* ratio increased dramatically during the boom—i.e., through 2006 —, the average loan-to-*value* (LTV) ratio showed little movement until 2006 (Gelain, Lansing, and Mendicino 2012). Hence the LTI seems to be a good indicator of the decline in underwriting standards, a decline that appears also if we regress the probability of approval on a range of observable characteristics, including the LTI, of borrowers and houses. In such regressions, the applicant's LTI is a weaker predictor of approval rates in 2006 than in 2000 (Dell'Arriccia, Igan, and Laeven 2009).

Approval rates for mortgages increased significantly during the boom: from 70.1% to 84.2% (Garriga 2009). Approval rates increased by 10.83 percentage points for white borrowers, 13.08 percentage points for black borrowers, and 12.6 percentage points for Hispanic borrowers. Moreover, origination rates (i.e., the percentage of applications that lead to a mortgage origination) increased from 53.1% to 66.2% from 2000 to 2006, which may be indicative of looser lending standards (Dell'Arriccia, Igan, and Laeven 2009) or of changes in the demand for housing or for credit. In section 4, we shall use an IV strategy to disentangle these determinants of the change in credit market equilibrium conditions.

This paper relates metropolitan area credit conditions to changes in segregation. Measuring lending standards at the metropolitan level makes sense as metropolitan areas are likely to be relevant credit markets. DiSalvo (1999) describes how US Federal Reserve branches use borders for credit market areas that are very similar to those of metropolitan areas. The Federal Reserve

[12] It is noteworthy that we use a measure of the *bank's* decision: approval or denial. A given denial rate can be associated with different origination rates, as households may withdraw their application from a particular bank; this explains our focus on approval rates.

[13] Descriptive statistics are presented in the appendix (Table A2).

banks in Dallas and Philadelphia use metropolitan statistical area (MSA) borders, and the other branches use borders that are close to these such as Ranally Metropolitan Areas.

3.4 OLS Estimates

At the state-level, Figure 1(i) displays a positive correlation between white isolation across urban census tracts and the growth of mortgage approval rates. Increases in black, white, and Hispanic segregation at the metropolitan-area level are also positively correlated with 2000–2006 increases in approval rates[14] (as well as with increases in the median LTI ratio).

Table 1 presents the results of two series of metropolitan-area level OLS regressions relating changes in segregation to changes in the metropolitan area's approval rate and also to changes in its median LTI ratio. The regressions are as follows:

$$\begin{aligned} \Delta Segregation_k &= \gamma_{Approval} \Delta Approval\ Rate_k + \alpha_{Approval} \Delta Demographics_k \\ &\quad + X_k \beta + State_{s(k)} + \varepsilon_k \end{aligned} \tag{1}$$

$$\begin{aligned} \Delta Segregation_k &= \gamma_{LTI} \Delta Loan - to - Income_k + \alpha_{LTI} \Delta Demographics_k \\ &\quad + X_k \beta + State_{s(k)} + \varepsilon_k \end{aligned} \tag{2}$$

where k indexes metropolitan areas (MSAs or μSAs) and $s(k)$ is the state of metropolitan area k. We use $\Delta\text{Segregation}_k$ to denote the 2000-2006 change in segregation in metropolitan area k, where segregation is measured in terms of isolation (columns 1–3) and exposure (columns 4 and 5) as defined in Section 3.2 for Blacks, Whites, and Hispanics. The term X_k is a set of observable controls, and $\text{State}_{s(k)}$ is a state fixed effect. The ε_k term is the residual clustered at the state level. Finally, $\Delta\text{Demographics}_k$ is a set of controls for the change in the fraction of Blacks, Hispanics, Asians, and other races in the metropolitan area.

Demographic controls capture part of migrations' impact on segregation, but such controls have surprisingly little effect on the coefficients for changes in approval rates and for changes in the LTI ratio. This result is well-illustrated on Figure 1(ii). Figure 1(ii) shows states with a large increase in Hispanic population (above the median increase across states) as blue dots and shows states with a small increase in Hispanic population as red dots. Regardless of which subset of states we focus on, the linear positive correlation between black isolation and changes in

[14] To clarify the exposition, we define the approval rate as "100% minus the denial rate". Thus, since relaxed lending standards lead to higher loan/income thresholds, they also lead to higher approval rates.

approval rates continues to hold.[15]

The OLS estimates γ_{LTI} and γ_{Approval} in metropolitan regressions cannot be interpreted as supporting causality; rather, they provide evidence of an economically and statistically significant correlation between changes in *equilibrium* credit conditions and changes in segregation. We investigate the possibility of a causal interpretation in Section 4.

Table 1, which reports the OLS results, is divided into three panels corresponding to black segregation (upper panel), white segregation (middle panel), and Hispanic segregation (bottom panels).

The estimates reported in column 1 of Table 1 (upper panel) suggest a positive and significant correlation between the increase in black isolation and either the increase in approval rate or the increase in LTI, when we control for changes in metropolitan racial demographics. The magnitude of these correlations are both reasonable and economically significant: a 13 percentage point increase in approval rates – the magnitude of the 2000–2006 change – is correlated with a 1.4 percentage point increase in black isolation. An increase of 0.4 in the median LTI ratio (i.e., the magnitude observed during the boom) is correlated with a 2.8 percentage point increase in black isolation. Our regression controls for a state effect that captures state-level unobservables as well as for demographic controls. Column 2 of Table 1 introduces both state effects and a set of additional variable controlling for mortgage credit risk[16], and for metropolitan area housing supply elasticity (see Saiz (2010)). These additional controls yield almost no change in the correlation between changes in lending standards and change in black isolation. Using the same specification, Columns 3 and 4 show that the increase in black isolation is mostly accounted for by the negative correlation between relaxed credit

[15] The same is true for the positive linear relationship between White isolation and changes in approval rates.

[16] Mortgage credit risk is proxied by the fraction of past due loans in 2006 from the Mortgage Banker Association, the fraction of high risk loans in 2006 as predicted by 1995 HMDA denial standards, the median mortgage spread over treasuries in 2000 and 2006, and the growth in missing income loan originations. HMDA glossary states that a mortgage for a single-family owner-occupied house has missing income information when an originating institution does not rely on the applicant's income. The fraction of high risk loans is constructed as follows: we first estimate how observable borrower, housing, and neighborhood characteristics predict approval probabilities in 1995. Using the coefficients estimated in this first step, in each metropolitan area, we estimate in 2000 and in 2006 the fraction of loans that would have had a low approval rate in 1995 (below the 25th percentile in 1995). This procedure yields an estimate of the fraction of high risk loans in 2000 and 2006, which we include, as control, in the regression.

standards and the exposure of Blacks to Whites.

Column 3 adds controls for price and income changes. Because some share of the LTI increase might be due to banks adjusting their lending standards to increases in house prices and/or to changes in households' income, in this column we incorporate not only the 2000–2006 log increase in the Case–Shiller index but also the log increase in the personal income of Whites, Blacks, Hispanics, and Asians for this period as computed using data from the American Community Survey. Yet the coefficients in column 3 remain positive and significant.

The middle and bottom panel of Table 1 presents similar regressions for Hispanics and Whites, respectively. Increases in both LTI ratios and approval rates are correlated with increases in white isolation (+2.4 for approval rates); these relaxed lending standards are also correlated with declines in the exposure of Whites to Blacks (−3.8 for approval rates and −1.5 for LTI ratios) but not significantly with increases in the exposure of Hispanics to Whites.

Confounding Factors

OLS results of Table 1 regress changes in segregation on changes in the median LTI ratio and approval rate to changes in measures of segregation. However, both approval rates and LTIs are equilibrium quantities – that is, they are determined at the equilibrium of mortgage credit markets. In Table 1, Ordinary least-squares (OLS) regression estimates control for income and demographic changes, two major determinants of housing and credit demand. However, other potentially unobservable demand shifters may be correlated with credit supply and confound our causal estimates of the impact of credit standards on metropolitan segregation. The key challenge of this paper is therefore to find an identification strategy that separates the effect of a rising credit supply from the effect of changes in the demand for credit.

We should all the more concerned about unobservable confounding factors that *observable* confounding factors – income and demographics – are correlated with the change in loan-to-income ratio and approval rate changes. Table A3, in the appendix, show that increases in loan-to-income ratios from 2000 to 2006 are positively correlated with same-period increases in wage income – on average and also for each racial and ethnic group evaluated (i.e., Blacks, Whites, and Hispanics).[17] The positive correlation we find between higher incomes and higher LTI ratios suggests instead the presence of upward housing demand shocks.[18] By contrast, changes in wage income appear uncorrelated with changes in approval rates, at the exception of change in whites wage income that display a small but significant negative relationship with approval rate changes.[19]

Demographic changes can also lead to shifts in mortgage credit demand. Table A3 in the appendix shows that demographic changes are correlated with approval rate changes but not with LTI changes. Approval rate changes are positively correlated with white inflows, and negatively correlated with minority inflows. Hence the specific role of relaxed lending standards might be confounded with demographic effects.

Altogether results of Table A3 suggest that, at a very minimum, an appropriate instrument for changes in lending standards should not display a significant correlation with income changes and demographic changes.

4. IDENTIFICATION STRATEGY AND RESULTS

4.1 Banks' Liquidity Levels and the Supply Channel of Credit Standard Relaxation

In this section we describe an IV strategy for the identification of the causal impact of relaxed credit standards on segregation. The instrumental variables are measures of banks' liquidity conditions at the metropolitan area level in the early 1990s – that is, before the mortgage credit boom (2000–2006) and also before a number of key transformations that affected the mortgage industry and favored the rise of securitization. The underlying hypothesis is that increased securitization allowed banks with initially low levels of liquidity to *catch up* by increasing their approval rate and median loan-to-income ratios relative to banks with initially high levels of liquidity. Securitization weakened the dependence of mortgage credit supply on local banking conditions as measured by bank liquidity.

---

[17] Changes in wage income are computed as changes in median wage income. Median wage income are calculated at the metropolitan area level using the 2000 and 2006 waves of the American Community Survey, from which annual measures are available for the largest 238 metropolitan areas.

[18] Ferreira and Gyourko (2011) indeed suggested that income shocks preceded house price increases at the census tract level

[19] This negative relationship is consistent with the evidence of Mian and Sufi (2011) of a negative relationship between changes in income and changes in credit, which they interpret as evidence of a credit supply shock.

The first such transformation favoring the rise in securitization of the industry was the Federal Housing Enterprises Financial Safety and Soundness Act of 1992.[20] This legislation established housing goals for the Government Sponsored Enterprises (thereafter GSEs, which include, but are not restricted to, Fannie Mae and Freddie Mac) regarding low- and middle-income applicants and previously underserved areas, and at the same time it gave preferential capital treatment both to the GSEs and to banks holding the mortgage-backed securities (MBS) issued by GSEs.[21] This reduced "capital charges" treatment for the GSEs overturned previous recommendations from the US Treasury that GSEs should actually increase their capital in order to comply with the Basel Committee's risk-based capital rules.[22] The second transformation involved the rapid development in the 2000s of an alternative securitization chain that packaged mortgage loan originations through asset-backed commercial paper conduits (ABCP) and that provided loan "warehousing" and ultimately securitization via private entities.[23] The development of ABCP conduits was boosted by new accounting rules,[24] which allowed assets in ABCP loan programs to be excluded from the risk-weighted asset base of sponsoring banks and so resulted in de facto regulatory arbitrage.[25] The ABCP market grew from about $600 billion in 2000 to $1.2 trillion in 2006, by which time it had become the largest US money market instrument. Moreover, the share of mortgage loans in new ABCP issuances increased from 36% in 2000 to 70% in 2006.[26] These two structural transformations were the main contributors to the mortgage industry's change from a predominantly "originate and hold" model to a

[20] The FHEFSA Act was not implemented immediately. The Act sets a deadline of December 1994 for GSEs to meet minimum capital requirements (Office of Federal Housing Enterprise Oversight 1998). With regard to the indicated special lending areas, HUD issued formal goals only in December 1995, following some interim goals for the period 1993-1995. In fact, the first major GSE announcement concerning these lending goals was made in 1994, when Fannie Mae made a $1 trillion commitment to affordable housing – which included money lent under less stringent underwriting standards. We assume that the FHEFSA act was fully implemented in 1995 and use information for the period 1990-1994 to construct our instrument.

[21] The GSEs were required to hold a capital buffer of only 0.45% as a guarantee against the default risk of the MBS they issued and only a 2.5% capital buffer against mortgages held on their own balance sheets; in comparison, federally insured banks were required to maintain a 4% capital buffer against their mortgage holdings. Furthermore, banks were required to hold only a 1.6% capital buffer against their holdings of MBS issued by the GSEs (Acharya, Richardson, and Van Nieuwerburgh 2011).

[22] US Treasury (1990)

[23] Levitin and Wachter (2010).

[24] These accounting rules were enacted following the 1992 Basel accord.

[25] Acharya, Schnabl, and Suarez (2010).

[26] Adrian and Shin (2010).

predominantly "originate to distribute" model. While the outstanding stock of home mortgages increased between 1990 and 2007 from \$2.52 trillion to \$11 trillion, the share of home mortgages securitized by GSEs or so-called private-label securitizers rose from 37% to 58.7%.[27]

Following Loutskina and Strahan (2009) and Loutskina (2011), we focus on the effect of bank liquidity – as measured by a bank's share of liquid assets or the securitizability of its loan portfolio – as a predictor of lending decisions: approval rates and loan-to-income ratios. We construct metropolitan area liquidity measures as follows. First we match individual mortgage originations (as reported in HMDA data) to the originating bank; the matching procedure is restricted to the sample of individual mortgages originated by banks reporting to the Federal Reserve, the Federal Deposit Insurance Corporation, or the Office of Comptroller of the Currency. Then we consider two liquidity measures: the share of liquid assets in total assets and the securitizability of portfolios. At the bank level, we follow Loutskina and Strahan (2009) and construct the first measure, bank-level liquidity, as:

$$Liquidity_{b,t} = \frac{Cash + Securities\ at\ time\ t\ in\ assets\ of\ bank\ b}{Total\ assets\ of\ bank\ b}$$

The second measure, bank-level securitizability, is constructed, following Loutskina (2011), as

$$Securitizability_{b,t} = \sum_{j=1}^{6} \frac{Securitized_{j,t}}{Loans_{j,t}} \cdot Share_{j,b,t}$$

where j indexes the type of loans in bank portfolios, $Share_{j,b,t}$ is the share of type j loans in bank b portfolio in year t, $Securitized_{j,t}$ is the economy-wide volume in USD of securitized loans of type j in year t, and $Loans_{j,t}$ is the total economy-wide loan volume in USD of type j in year t.[28] This latter measure can be viewed as a weighted average of the potential to securitize loans of a given type (based on market wide averages), where the weights reflect the composition of an individual bank's loan portfolio.

The corresponding metropolitan area measures are the *average* of each bank's liquid assets ratio and securitizability index weighted by the volume of originations (measured in US dollars) in the

[27] These figures are derived from the US Flow of Funds accounts, table L.218.

[28] Loan portfolios are broken down into six different types of loans (Loutskina, 2011): (i) home mortgages, (ii) multi-family residential mortgages, (iii) commercial mortgages, (iv) consumer credit, (v) business loans not secured by real estate (commercial and industrial loans), and (vi) farm mortgages. Securitizability measures are based on the US Flow of Funds accounts, and individual bank-level loan data are from each bank's Report of Income and Condition.

area:

$$Liquidity_{k,t} = \sum_{b=1}^{B_k} Fraction_{b,k,t} \cdot Liquidity_{b,t}$$
$$Securitizability_{k,t} = \sum_{b=1}^{B_k} Fraction_{b,k,t} \cdot Securitizability_{b,t}$$

where b sums over the B_k banks that originated mortgages in metropolitan area k. $Fraction_{b,k,t}$ is the fraction of mortgages originated by bank b in metropolitan area k in year t. These measures are computed for each year between 1990 and 1994.

Finally, we average the metropolitan area liquid asset ratio (resp., securitizability) over the 1990–1994 period and use them as instruments for the 2000–2006 growth in the median LTI ratios (resp., the approval rate). There is significant independent variation of our liquidity and securitizability measures as almost two thirds of the variation in liquidity ratios across metropolitan areas is not explained by the securitizability measure. By averaging our liquidity measure over five years (1990-1994) we screen out the effects of year-to-year variations in banks' balance sheet on liquidity.

Our hypothesis is that lending standards in a metropolitan area are correlated with the liquidity position of banks active in that area but that the correlation is weakened by rapid development of securitization, which leads to a growing disconnect between liquidity and lending standards. As a consequence, metropolitan areas with an *initially low level of bank liquidity* should experience a *greater relaxation of their lending standards* than do metropolitan areas with an initially high level of liquidity.[29] Figures A1 and A2 in the appendix confirm that the average 1990–1994 metropolitan area liquidity measures are positively correlated with the metropolitan area approval rate's *level* in 2000. Figure 2 (top panels) shows that low-liquidity and low-securitizability banks "caught up," i.e. there is a negative correlation between our initial liquidity measures and the *change* in lending standards from 2000 to 2006. Figure 2(i) shows that metropolitan areas with initially low levels of securitizability experienced a greater increase in

[29] Note that while we are using the same liquidity measure to identify credit-supply related shocks, our instrumental variable approach differs from that of Loutskina and Strahan (2009) in some important dimension. Loutskina and Strahan (2009) analyze how *current* bank liquidity conditions affect the approval of jumbos loans at the *bank level*. By contrast, we exploit the time series dimension of the data and study how banks' liquidity measures aggregated at the *metropolitan area level* in 1990-1994 predict the change in lending standards between 2000 and 2006 at the *metropolitan area level*.

approval rates during the boom than did metropolitan areas with initially high levels of securitizability. Figure 2(ii) shows similar (albeit less strong) results regarding the link between bank liquidity and LTI ratios.

LTI growth and approval rate increases are indeed significantly and positively correlated with the growth of securitization volumes. Figure 2 (bottom panels) plots the growth in the volume of securitizations across metropolitan areas ranked according to their initial securitizability index (panel iii) or their initial liquid asset ratios (panel iv). The figure illustrates that both low- and high-liquidity approval rates experienced a securitization boom during the period 1995–2006 but that the boom was more pronounced for low-liquidity metropolitan areas.

This is confirmed in first-stage regressions of approval rate changes (resp. LTI changes) on metropolitan area securitizability measures (resp. liquidity measure), which are presented in Table 2 (resp. Table 3). Column 1 of each table presents the regression with demographic controls and no state effects. The coefficient is significant at 1% in both tables, and the *F*-statistic for the securitizability (resp. liquidity) coefficient is 13.7 (resp. 8.5). Although the *F*-statistic for approval rate changes is strong, the *F*-statistic for LTI changes may be considered weak for the LTI IV regression. Therefore, we later check our corresponding Cragg–Donald statistic to get an estimate of the relative size of the finite sample IV bias.[30] First-stage regressions of Tables 2 and 3 explain from 4.3% to 17.5% of the dependent variable's total variance. Column 2 of each table adds state effects, so that the impact of liquidity and securitizability is estimated within each state.[31] The coefficients are significant at 1% and are not statistically different from those reported in column 1, indicating that first-stage results are robust to the inclusion of state effects. The last column of each table regresses the log change in the volume of securitization on each of the instrument, confirming that metropolitan areas with low level of liquidity or securitizability experienced a higher growth of securitization during the boom.

Three potential concerns affect the validity of our instrument. The first concern is that banks' liquidity and loan portfolio securitizability in 1990-1994 may reflect fundamental demand characteristics rather than being determinants of future credit supply shifts. In particular, loan

[30] Generally speaking, IV estimators are asymptotically consistent but biased in finite samples; the Cragg-Donald test statistic provides an estimate of the maximum size of such finite-sample bias. Results are robust to such corrections.

[31] Throughout the paper, panel data regressions with fixed effects are estimated using the within estimator. Using OLS with dummy variables does not significantly affect estimates.

portfolio securitizability may simply reflect the bank's specialization in mortgage credit (as opposed to other kinds of credit such as consumer credit or business loans), which may be related to non–time-varying consumer preferences for homeownership; if so, then consumer demand rather than credit supply would explain the correlation between loan portfolio securitizability and the 2000–2006 growth in approval rates and LTI ratios. We address this issue by rerunning the first-stage regression while using two alternative measures of liquidity and securitizability. First, we focus on the liquidity of banks whose share of originations in that metropolitan area accounts for less than 25% of all its originations. For these banks, the specific metropolitan area represents such a minority share of their holdings that metropolitan area demand is unlikely to have much effect on their balance sheet liquidity and loan portfolio composition. The results based on this measure are presented in column 3 of Tables 2 and 3. The coefficients are similar and not statistically different from those in column 1 of these tables. Second, we focus on liquidity and loan portfolio securitizability for banks' balance sheets while *excluding* residential mortgages; this measure is less likely to reflect households' preferences for homeownership or credit.[32] We find that the regressions using this alternative measure of securitizability yield results that do not differ significantly from those when residential mortgages are included.

The second concern is that our instrument maybe correlated with potential demand shifters proxied by income changes and demographic changes. In order to assess the magnitude of these issues, we recompute the correlations of Table A3. This time, instead of correlating lending standards with observable demand shifters, the table correlates these demand shifters with our instruments. The results, presented in Table A4.1, indicate that none of our instruments display any significant correlation, with changes in wage income, or demographic changes. While such lack of correlation is not a direct proof of exogeneity, it is strongly reassuring given that non-instrumented changes in the loan-to-income ratio and approval rate are correlated with several of these observable factors.

The third concern is that our instruments might be correlated with other structural determinants of racial segregation that could vary over time *or* whose impact on segregation

[32] These alternative measures are computed in the following way. Loan portfolio securitizability is the securitizability measure computed on other loans than mortgages. The liquidity measure excluding mortgages is cash+securities divided by total assets excluding mortgages.

varies over time. Table A4.2, column 3, shows indeed that the change in income inequality (Gini 2010 minus Gini 2000), the burglary rate, the homicide rate, the college premium, and the average July temperature are for the most part positively correlated with metropolitan black isolation in 2000. Formally column 3 displays the coefficient of the regression of these measures of metropolitan amenities on black isolation in 2000. Literature in sociology has suggested that racial segregation and crime are strongly related, with the direction of causality likely going both ways (Peterson and Krivo, 1993). The lack of significant correlation of our instruments with measures of overall income inequality, or of racial inequality (the blacks/whites income gap) is important. It implies, in particular, that the finding that that reduced racial inequality may lead to increased racial segregation (Bayer, Fang, McMillan, 2005), is unlikely to drive our IV estimation results. In contrast to the correlations of column 3, column 1 (resp., column 2) shows the coefficients of the regressions of the same potential confounders on our liquidity instrument (resp., our securitizability instrument). The columns suggest that there is no evidence of a statistically significant correlation of the instruments (liquidity and securitizability) with such confounders.

4.2 Instrumental Variable Estimates

Our IV estimates are presented in Table 4. The change in loan-to-income ratio is instrumented by the metropolitan area bank liquidity level between 1990 and 1994. Approval rate changes are similarly instrumented by the metropolitan area bank securitizability index. Diagnostic statistics for the instrumental variables estimation are presented in the last two rows of each panel. For the approval rate regressions, the first stage *F*-statistic has a *p*-value below 1% and significantly above 10. The Cragg–Donald *F*-statistic is substantially above 16.38, the critical value for a maximum 10% finite-sample bias. For the LTI ratio regressions, the Cragg–Donald *F*-statistic is substantially above the critical value for a maximum 15% bias and close to the critical value for a maximum 10% bias. Given the size of the OLS and IV estimates, a 10%-15% bias would not significantly alter this paper's findings. For the LTI ratio regression, the first stage *F*-statistic also has a *p*-value below 1% (0.0053) although the statistic is around 8.5. Overall, those statistics do not present significant evidence that instrument weakness would substantially alter our main conclusions.

In Table 4, the impact of the approval rate on black isolation is significant, strong, and positive. A one–standard deviation increase in the approval rate, i.e. a 10 percentage point increase, is associated with a 1.8 percentage point increase in black isolation, mostly due to a decline (also of 1.8 percentage points) in the exposure of Blacks to Whites; in this case there is also a (nonsignificant) decline of 0.1 percentage points in the exposure of Blacks to Hispanics.

Higher LTI ratios had a similar positive effect on black isolation (second panel of Table 4). An increase in the LTI ratio by 0.4 (the national average increase) leads to an 8.1 percentage point increase in black isolation (20.249×0.4). This finding reflects the combination of a lower exposure of Blacks to Hispanics and of Blacks to Whites. The latter effect explains about 60% of the overall positive effect of rising LTI on black isolation. Reduced form coefficients (presented in column 4 of table 4) also suggest that metropolitan areas where banks were more liquid and had more securitizable loans in 1990–1994 witnessed a significantly greater 2000–2010 decline in black isolation.

On average, black exposure to white neighbors *declined* from 33.2% percent in 2000 to 32.7 percent in 2010 (–0.5 percentage points) in the average metropolitan area (Table A1 of the appendix). Therefore, the numbers reported here suggest how overall changes in mortgage lending standards have contributed to lowering the exposure of Blacks to white neighbors. A simple accounting exercise using the IV estimates suggests that black exposure to Whites would have been 2.3 percentage points higher (resp., 5.2 percentage points higher) had approval rates (resp., LTI ratios) stayed constant from 2000 to 2006. These are sizable impacts, of about 45% of (resp., 1.1 times) the average decennial change in black isolation (-5.1 percentage points).

The OLS estimates are substantially lower than the IV estimates. For instance, OLS (resp. IV) regressions with state effects and controls predict an increase of 1.2 percentage points (resp. 1.8 percentage points) in black isolation for a 10 percentage point increase in the approval rate. This difference suggests that unobservable upward demand shocks tend to lower black isolation. We indeed observe that rising LTI ratios are more strongly correlated with black wage income growth than with Hispanic and white income growth. This suggests that, in OLS results, upward black housing demand shocks confound OLS estimates; as such OLS point estimates underestimate the segregative impact of the relaxation of lending standards.

When instrumenting credit standards, the overall effect on Hispanic segregation is nonsignificant. This result is not likely to be the outcome of a weak instrument, given that we have reported high values for both the Cragg–Donald and the first stage F-statistic. However, the statistical nonsignificance of this effect masks two underlying trends of interest: higher LTI ratios tend to *increase* the exposure of Hispanics to Whites (by 4.1 percentage points for a 0.4 increase in LTI) and also to *reduce*, albeit nonsignificantly, the exposure of Hispanics to Blacks.

LTI increases have a significant negative impact on the exposure of Blacks to Hispanics but only a mild one on the exposure of Hispanics to Blacks; these findings are consistent with evidence that Hispanic population increased substantially in some metropolitan areas while black population increased by smaller numbers. Such results suggest that new Hispanic households typically settled into either Hispanic enclaves or black neighborhoods, and thus black exposure to Hispanics changed more significantly than Hispanic exposure to Blacks.

Endogeneity of Migration Flows

In our OLS and IV regressions, covariates are included for the change in the fraction of Blacks, Hispanics, and Asians in the metropolitan area. However, an argument could be made that these net changes in racial demographics are endogenous to the metropolitan area supply of credit.

Suppose, for example, that an increased credit supply facilitates Hispanic mobility to metropolitan areas with large Hispanic enclaves. Then metropolitan areas that increase their supply of credit see larger inflows of Hispanics into Hispanic enclaves, and so metropolitan areas with more relaxed lending standards will tend to show stronger effects of net Hispanic population flows on both black isolation and white isolation. We circumvent this identification problem by using the Hispanic fraction of the metropolitan area in 1980 to predict the increase in that fraction while using the same (2003) metropolitan area borders throughout the analysis. Indeed, evidence from prior literature suggests that Hispanics are more likely to move to metropolitan areas that already contain large Hispanic enclaves (Card 2009). Those metropolitan areas with a substantial fraction of Hispanics in 1980 witnessed greater increases (than did other metropolitan areas) in that fraction: a metropolitan area with 10% more Hispanics in 1980 saw a 0.4 percentage point greater increase in that percentage over the period 2000–2010. The *F*-statistic of the simple OLS regression is 40.7.

Table A5 in the appendix replicates our main IV regressions when the dependent variable is black isolation or the exposure of Blacks to Whites or Hispanics and after replacing "2000–2010 change in the metropolitan area's Hispanic fraction" with "change in Hispanic population as predicted by the metropolitan area's 1980 Hispanic fraction". The coefficients of interest remain positive and strongly significant.

The role of housing supply elasticity

While up to Section 4 the paper presents results at the metropolitan level, Section 5 will present neighborhood-level, i.e. census tract level, evidence on population flows and prices that shed light on possible mechanisms underlying these IV results. In particular, tract-level findings will illustrate the role of general equilibrium effects of lending standards on prices: house prices in increased significantly more in mostly white neighborhoods than in mostly blacks or racially mixed neighborhoods.

At the metro-area level, Glaeser, Gyourko, and Saiz (2008) show that price increases during the last credit boom have been much larger in metropolitan areas with low housing supply elasticity. Such price increases have been larger in metro areas with low housing supply elasticity and declining lending standards (Imbs and Favara, 2013). Saiz (2010) measures housing supply elasticity using the amount of developable land in U.S. metropolitan areas.

In highly inelastic metropolitan areas, more desirable neighborhoods have witnessed larger house price increases than less desirable neighborhoods. Indeed, during the boom of 2000—2006, the metropolitan area standard deviation of tract-level price changes was positively correlated with metropolitan area average price increases. Also, more inelastic metropolitan areas saw greater average price increases and a greater *standard deviation* of price increases within each metropolitan area.[33]

We test whether the segregative impacts of the credit boom on segregation are stronger in more inelastic metropolitan areas. The underlying hypothesis is that, in desirable neighborhoods, credit-driven changes in housing demand are more likely to translate into price increases than into new construction.

Table A6, in the appendix, presents IV estimates in which changes in lending standards are interacted with the metropolitan areas measures of housing supply elasticity estimated by Saiz

[33] The MSA-level correlation between the 2000-2010 log increase in house prices and the 2000-2010 log increase in the standard deviation of house prices is 0.77 and significant at 1%. The correlation between housing supply elasticity and the log increase in the standard deviation of house prices is –0.22 and significant at 1%. Section 5 describes the data sources for census-tract level house prices.

(2010).[34] The upper panel presents the results for approval rates while the bottom panel presents results for the LTI. In both panels, the interaction between housing elasticity and lending standards is negative and significant for black isolation, and positive and significant for the exposure of Blacks to Whites. The non-interacted coefficients for changes in lending standards keep the same sign as in our baseline IV estimates and remain significant. These regressions results show that the effects of lending standards on segregation critically depend on the degree of housing supply elasticity. For the metropolitan with average elasticity (2.49), a 10 ppt increase in the approval rate leads to a 0.86 ppt increase in black isolation (first line, upper panel), and a 0.4 increase in the median loan-to-income ratio leads to a 11.74 x 0.4 = 4.7 ppt increase in black isolation. In metropolitan areas with low housing supply elasticity, e.g. the 25th percentile of housing supply elasticity 1.65, the negative effect of lending standard relaxation is stronger: a 10 ppt increase in approval rate leads to a 0.86 – 0.10 x 4.8 x (1.65-2.49)= 1.26 ppt increase in black isolation, an effect that is 47% larger than for the average metropolitan area. In contrast, in metropolitan areas with high housing supply elasticity (e.g. the 75th percentile of housing supply elasticity 3.09) the effect is smaller and not significant: a 10 ppt increase in approval rate leads to a 0.86 – 0.10 x 4.8 x (3.09-2.49) = 0.57 ppt increase in black isolation.[35] Similar results apply for the impact of loan-to-income ratio increases (lower panel).

4.3 Discussion: The Role of the Foreclosure Crisis

An important issue deals with the timing of the measurement of changes in segregation. Tract-level US Census data is decennial, hence changes in urban segregation are calculated over the 2000–2010 decade, while credit measures are measured between 2000 and 2006. Years following the 2000–2006 boom could be a confounding factor because post-boom events between 2007 and 2010 – including foreclosures, labor market shocks, and other factors – may

[34] The interaction of lending standards with housing supply elasticity is instrumented by the interaction of liquidity measures with housing supply elasticity. Note that housing supply elasticity is not significantly correlated with liquidity once state effects are controlled for.

[35] The results for approval rates and LTI are consistent. Note however that in these regressions, with two instruments and two instrumented variables (lending standards and lending standards interacted with housing supply elasticity), the instruments are valid in the case of approval rates (with a value of about 8 for the Angrist-Pischke F statistic for the main instrument and 29 for Cragg Donald implying a maximum bias of less than 5%) but weak in the case of LTI (with a value of about 2.6 for the Angrist-Pischke F statistic for the main instrument and 4.1 for Cragg Donald implying a maximum bias of 25 %).

be correlated with changes in credit conditions during the boom years. We tackle this issue in two ways: first by using foreclosure data to control directly for the effect of the foreclosure crisis; second by using two additional datasets where the timing of demographic data more closely coincides the timing of credit data.

Controlling for Foreclosures after the Credit Boom

Omitting the impact of the foreclosure crisis in our main regressions could have led to either an overestimation or an underestimation of the impact of lending standards. First, to the extent that household mobility to same-race neighborhoods was reduced by the foreclosure crisis that began in 2007, as foreclosure rates were higher in metropolitan areas with a larger decline in lending standards, our results *under*estimate the credit boom's effect on racial segregation. Second, foreclosures could have exacerbated racial segregation if households most affected by foreclosure sales were households who had joined mixed neighborhoods during the credit boom – or simply if the foreclosure crisis tended to affect black households relatively more than white and Hispanic households. So if foreclosures tend to magnify segregation, then our main estimates are upward biased and thus we have *over*estimated the boom's impact on racial segregation.

Our foreclosure data are from RealtyTrac and include the yearly number (by ZIP code) of notices of trustee sale and notices of foreclosure sales. We aggregate these numbers by metropolitan area to compute the number of foreclosures sales per 100 residents in each metropolitan area.[36] The median number of foreclosure rate per 100 residents is 0.4, and the standard deviation is 1.

We estimate whether foreclosure sales influence the estimates of our paper's main focus, the credit boom's effect on black isolation, by including that variable in the regression. This regression includes demographic controls, state effects, and metropolitan area elasticity of housing supply; as before, the standard errors are clustered by state.

[36] Our focus is on the sum of notices of trustee sales (NTS) and notices of foreclosure sales (NFS), which is the total number of foreclosure sales in the metropolitan area. We focus on actual sales because it may take more than a year for the foreclosure process to result in an eviction (as many as 455 days in the state of New York, for example).

Results for the approval rate regression are given in the upper panel of Table A7 in the appendix, and results for the LTI regression are given in the lower panel. Overall, the coefficients of interest remain significant, strong, and of the same sign as the original OLS and IV regressions. We present in Table A7 OLS rather than IV regressions. Indeed, without an additional instrument for the foreclosure sales rate this variable will be in effect instrumenting the growth of lending standards (LTI and approval rates). Reassuringly however, repeating our baseline IV estimation (Table 3) with the foreclosure sales as an additional control yields very similar estimates.[37]

In the approval rate regression (Table A7, upper panel), the coefficients for foreclosure sales is only significant in column 3 where an increase of 1 foreclosure sale per 100 residents (which corresponds to a one-standard deviation change) lowers the exposure of Blacks to Hispanics by 0.207 percentage points. Introducing foreclosure sales as additional control leaves almost unchanged the coefficients of the effect an increase in approval rates on black isolation, exposure of Blacks to either Whites or Hispanics (Table 1, col.2, col.4, col.5). In the LTI regression (Table A7, lower panel), there is a (weakly) significant effect of foreclosure sales on segregation: an increase of 1 foreclosure sale per 100 residents reduces black isolation by 0.49 percent. Here again, controlling for foreclosure sales leaves the effect of an increase of LTI on black segregation virtually unchanged. In sum, the effect of foreclosures during the bust appears largely orthogonal to the segregative effect of a relaxation of lending standards. However a small share of that segregation increase has been "unwound" by foreclosure sales.

Estimates from American Community Survey Data.

The 2006-2010 American Community Survey (ACS), made available in 2010, provides census-tract level racial demographics. Because the survey collected information from 2006 to 2010, the median household was surveyed in 2008, 2 years earlier than the 2010 decennial census. Also, according to data collected by RealtyTrac, national foreclosure volumes started to increase around mid-2007, and thus 2008 ACS demographics are arguably less affected by the foreclosure crisis than those of 2010 Census.

Table A8 presents results from OLS regressions with demographic controls and state

[37] Estimations results available upon request.

effects (Column 1), with additional controls for proxies of applicants' creditworthiness (Column 2), and with income and price controls (Column 3). Columns 4 and 5 present results on black exposure to Whites. Column 6 presents the IV isolation for black isolation. Results are robust and similar to the main baseline OLS and IV estimates: lending standards led to an increase in black isolation, as measured in the ACS.

Estimates using School Segregation data

We can also make a step toward resolving the same issue of events in 2007-2010 potentially confounding estimates through the use of school-level student demographics, which are available *annually* from the US Department of Education's Common Core of Data for all US schools (Public and Private School Universe). Hence school data is available for the same period as the credit boom, from 2000 to 2006. Here we focus on primary schools, but our results are robust to using secondary-school segregation. School segregation is of interest per se as some school peer effects have been shown to have a causal impact on students' educational achievement (Epple and Romano 2011).

We estimate the effect of the 2000–2006 changes in credit conditions on the change in segregation across schools for the same period. School segregation measures are computed in the same way as urban segregation measures, by using school-level student demographics instead of census-tract level resident demographics. Hence the isolation of black students is the average fraction of black peers for an average black student.

Table A9, upper panel, presents the regression results based on school segregation when the relaxation of lending standards is measured by the change in approval rate. The first and the second column reproduce, for the case of school segregation, the OLS regression (Table 1, col.1 and col.2, upper panel), and the third column the IV regression (Table 3, col.1, upper panel). The effect of the increase in approval rate on black students isolation are very similar (and not satistically different) to the effect on black residents isolation. The negative effect of the relaxation of lending standards on the exposure of blacks to whites is also very similar for school and residential segregation. Table A9, lower panel, show results for the change in LTI ratios; although point estimates are smaller than in the residential case – the link between residential and school segregation might have been altered by integration policies – the point estimate is not statistically different than for the effect of LTI on urban segregation. Overall the results on school segregation, for the period 2000-2006, are similar to those on urban segregation, for the period 2000-2010.

5 INSPECTING THE MECHANISM: MICRO EVIDENCE OF POPULATION FLOWS

5.1. Census Tract Demographic Changes

Results at the metropolitan area level suggest that the relaxation of lending standards led to an increase in black isolation, mostly because of a lower exposure of Blacks to Whites. However, lower black exposure to Whites could be due to a combination of (i) white households' mobility out of racially mixed neighborhoods and into predominantly white neighborhoods and/or (ii) black households' mobility out of racially mixed neighborhoods and into predominantly black neighborhoods.[38]

We tell these two different stories apart by build a neighborhood dataset: a matched longitudinal census tract level dataset based on the 2000–2010 US Census. We restrict our analysis to metropolitan areas with a black population of at least 100,000 individuals in order to focus on racial dynamics that could potentially be quantitatively significant.[39] The sample thus reduces to 64 metropolitan areas accounting for 75% of the total urban black population, and includes 40,050 census tracts.

We estimate census-tract level white and black demographic changes. Demographic change by race, in census tract *c*, is measured, following Card, Mas, and Rothstein (2008), as the change in white population as a fraction of the tract's total population in 2000:

$$White\ \ Population\ Change_c = \frac{Whites\ 2010_c - Whites\ 2000_c}{Population\ 2000_c}$$

Census tracts are grouped into $K=20$ bins (indexed by k), ranked according to their fraction of black residents in the 2000 Census: the first and 20th bin have, respectively, the lowest and highest average fraction of Blacks in 2000.

We measure the impact of the metropolitan area level relaxation of lending standards on

[38] These neighborhoods include desirable neighborhoods with a significant fraction of college educated black neighborhoods. Bayer et al. document the emergence of such neighborhoods between 1990 and 2000, in metropolitan areas with a growing fraction of college-educated blacks.

[39] Our results are however rather similar if we consider the entire set of metropolitan areas or if we use the threshold of Bayer et al. (2005): metro areas with population over 100,000 inhabitants and at least 10,000 blacks. The corresponding graphs are available from the authors upon request.

census tract demographic changes in the following way. First, we sort metropolitan areas in two groups, those metropolitan areas with a low *predicted* approval rate increase (lower than 13 percentage points), and metropolitan areas with a high *predicted* approval rate increase (higher than 13 percentage points). We use the predicted approval rate increase from the first stage regression of the metropolitan area IV results. In that regression, as in previous sections of the paper, the metropolitan area 2000-2006 approval rate increase is regressed on the approval rate instrument, namely the metropolitan area securitizability measure. Using the predicted approval rate instead of simply the approval rate increase substantially alleviates endogeneity issues surrounding the use of metropolitan area changes in approval rates. The dummy variables for the 20 black fraction bins are interacted with dummies indicating whether the census tract belongs to a metropolitan area for which the predicted approval rate change is low or high.

The regression, specified as:

$$\begin{aligned} White\ Pop.Change_c &= \sum_{k=1}^{K} (\delta_{k,high} Bin_k \times High\ Predicted\ \Delta Approval\ Rate_{metro(c)} \\ &+ \delta_{k,low} Bin_k \times Low\ Predicted\ \Delta Approval\ Rate_{metro(c)}) + \alpha \Delta Demographics_{metro(c)} \\ &+ State_{s(c)} + \varepsilon_c \quad (3), \end{aligned}$$

includes state fixed effects and metropolitan-level measures of demographic changes by race. The regression clusters standard errors at the state level. The state fixed effects have a zero average and the metro-area demographic changes are cross-sectionally demeaned, so that that the coefficients $\delta_{k,high}$ and $\delta_{k,low}$ can be interpreted as demographic changes for a tract in bin k of the average metropolitan area.

As in Card, Mas and Rothstein (2008), conditioning by metropolitan area demographic changes by race in the specification is key. These demographic variables do not only capture population flows, by race, into and out of the metro area. They also control for internal population dynamics due to birth rates/death rates differentials. Providing that these differentials vary between races and between metropolitan areas, but not by race across tracts within metropolitan areas,[40] the associated demographic changes, at the tract level, are indeed captured

[40] Specifically, using American Community Survey data for 2010, we find that the difference between black and white tract-level fertility rates is not significant at 95% when controlling for a metropolitan area

by the metro-level demographic controls by race. The coefficients of interest, $\delta_{k,high}$ and $\delta_{k,low}$, are thus capturing how *inflows and outflows* of populations into and out of the census tracts vary between tracts with high (low) predicted change in approval rates, for different shares of black population in 2000. Figure 3 illustrates regression results by plotting the estimated coefficients $\hat{\delta}_{k,high}$ and $\hat{\delta}_{k,low}$ against the share of black population in 2000, in the corresponding bin k. Star labels indicate whether the δ_k coefficients for high and low approval rate are statistically different from each other, thus showing whether patterns of inflows or outflows are significantly impacted by changes in approval rates.

Figure 3, panel (i) shows that white households moved in to census tracts with *less* than 7-8 percent of Blacks in 2000, but only in census tracts with a high predicted increase in approval rates. In those tracts with high predicted approval rate increase, white inflows range from about 9 percent (in almost completely white tracts) to 2 percent (in tracts with about 6 percent black population). In sharp contrast, in metropolitan areas with low predicted approval rate increase, white households did *not* move in to tracts with a similarly low fraction of black population. Racially mixed tracts – i.e. with an initial share of black population ranging from 10 percent to 60 percent – have all experienced white outflows. These outflows are about two to three times larger in tracts with high predicted increase in approval rates (ranging from 7 to 10 percent). Similar white outflows occurred in predominantly black tracts (tracts with 60 percent of more Blacks) at the exception of the very last bin of tracts, where there were almost no white resident in 2000.[41]

Put together these facts indicate a dynamic of white mobility *out of* black neighborhoods (more than 60% black), *out of* mixed neighborhoods (between 10% and 60% black) and *into* predominantly white neighborhoods (less than 10% of black population). These figures are consistent with the estimates of a tipping point model (Card, Mas, and Rothstein, 2008), which predicts significant white outflows from tracts that have between 5 and 20% of racial minorities – i.e. our estimate (around 10%) is within these bounds. The novelty of this paper's results lies in the fact that these dynamics of segregation have been considerably more pronounced in

fixed effect. The difference is of a small magnitude compared with the observed magnitude of population changes (the difference between a black woman's probability of having a child in the previous year and the same probability for white women is +0.1 percentage point).

[41] In that bin, tracts are about 95 percent blacks.

metropolitan areas that experienced a larger relaxation of lending standards during the credit boom.

Figure 3, panel (ii), describes black inflows for the same census tract bins as in panel (i). We observe sizeable inflows of Blacks into racially mixed neighborhoods (between 10% and 60% black), along with very large outflows of blacks from predominantly black neighborhoods (more than 60% black). These flow patterns are significantly more pronounced in areas with high predicted approval rate increases. Inflows peak at 7.2 (3.4) percent in tracts with 20 percent of black population and high (low) predicted increase in approval rates. In tracts with an 80 percent initial share of black population, outflows are about twice larger (11.1 vs. 5.8 percentage points) in census tracts located in metro areas with a high predicted approval rate increase.

Racially mixed neighborhoods (i.e. neighborhoods with 10-60% black population) experienced white *out*flows and black *in*flows over the same time period. Furthermore these black inflows and white outflows are of very similar magnitude, especially for neighborhoods with initially 15-20 percent of black population. This suggests that outgoing Whites have been replaced by incoming Blacks in those racially mixed neighborhoods, a pattern stronger in areas with high predicted approval rate increases. As a result these neighborhoods end up having a significantly higher share of Blacks in 2010 than in 2000, thus implying a reduction of the exposure of Blacks to Whites in those tracts.

Finally, the large outflows of Blacks from mostly black neighborhoods are driving a significant population decline in those mostly black tracts. As shown in results available from the author, these population declines have been facilitated by declines in lending standards: in tracts with 80 (95) percent of blacks, population has shrunk by about 10 (15) percent in areas with high predicted approval rate increases, but only by about one (10) percent in areas with low predicted approval rate increases.

To sum up, the findings just described allow an interpretation of our metro-level findings (Section 4) as the result of two mobility trends, which were both significantly more pronounced in metro-areas with a larger predicted relaxation of lending standards. The first trend is that of black mobility out of mostly black neighborhoods and into racially mixed neighborhoods. This should have contributed to an increased exposure of Blacks to Whites. But a second trend is that of white mobility out of racially mixed neighborhoods and into mostly white neighborhoods. This second trend has contributed to the reduction of the exposure of Blacks to Whites. This micro evidence provides an explanation for the metropolitan level findings: the relaxation of lending standards has led to a decline in the exposure of black households to white neighbors.

5.2 Explaining White and Black Mobility: Price Differentials and General Equilibrium Effects

Our tract-level results show that credit standards allowed significant black population moves from highly segregated neighborhoods into neighborhoods that were racially mixed in 2000. The relaxation of borrowing constraints should indeed give black households a greater choice set – i.e. a larger set of neighborhoods to choose from and the possibility of choosing homeownership over rental in more locations – at given prices, given neighborhood demographics, and given housing supply in each neighborhood. Black households could thus have taken advantage of greater credit availability to "trade up" by moving out of a segregated neighborhood and relocating into more racially mixed neighborhoods.

However, our results also suggest that lending standards did not facilitate black mobility towards mostly white neighborhoods (fraction black lower than 10%), and that, on the other hand, lending standards *did* make it easier for Whites to move to those mostly white neighborhoods.

One potential explanation may be that black households have a stronger preference for mixed neighborhoods than for mostly white neighborhoods. As Bayer, Fang and McMillan (2005) points out, black households face a trade-off between neighborhood quality and their preference for same-race neighbors, and racially mixed neighborhoods may achieve this balance between those two conflicting goals.

Another reason explaining the lack of black mobility towards mostly white neighborhoods is that the partial equilibrium view we outlined – lending standards expanding black households' choice set – abstracts from a general equilibrium effect of demand on prices. In particular, a greater credit supply leads to changes in house prices. The upward pressure on house prices is magnified in mostly white neighborhoods if white households' willingness-to-pay for local amenities or same-race neighbors is high but was previously constrained by credit availability. When credit gets relaxed, such households use credit to bid up prices, which prices out other groups, including black households. Evidence of the impact of credit supply on house prices is for instance presented in Imbs and Favara (2010), which showed that greater credit supply causes an increase in average MSA-level house prices.

We present evidence of this mechanism using census tract median house value data. Price data at the tract level is typically difficult to obtain for the entire universe of metropolitan areas. Detailed transaction-level data is usually available for subsamples such as the San Francisco Bay Area. At the national level, the 2000 long form version of the Census and the 2010 tract-level version of the American Community Survey provide self-declared estimates of house values that can allow us to build a 2000-2010 longitudinal sample of tracts. Goodman and Ittner (1992) and Bucks and Pence (2006) are somewhat reassuring about the quality of self-declared house values in that they conclude that, although such house value measures are noisy, measurement error is not significantly related to characteristics of the house, the owner, or the local market.

Figure 4 plots tract level log house value appreciation between 2000 and 2010 for the 20 bins of tracts sorted by the fraction of Blacks in the neighborhood. The graph was constructed as in the previous section presenting results on population flows, except that the dependent variable is the log house price appreciation. The regression also includes a state fixed effect, and a control for the 2000-2006 log change in the metropolitan Case Shiller index from Standard and Poor's. We also account for the 2007-2010 decline in house prices by including the 2007-2010 ZIP-level foreclosure rates per housing unit from RealtyTrac, although such controls do not significantly affect estimates once the 2000-2006 metropolitan area house price change is included. Each point on the graph corresponds to the price increase as predicted by the regression, not including, in other words "purged from the impact of", state effects, metropolitan demographic changes, metropolitan area log price increase, and the impact of foreclosures, but including the constant capturing a common national component (+42.7%)[42]

Figure 4 indicate that mostly white tracts (tracts with with fewer than 10% of black population) saw significantly higher price increases in metropolitan areas with a higher-than-the-median predicted approval rates (greater than 13 percentage points) than in metropolitan areas with lower-than-the-median approval rates (lower than 13 percentage points). The differences in price increases are both large – ranging between 6 and 18 percent – and for most of the bins of

[42]The inclusion of the constant is the prediction does not obviously affect the differences between the points on the graphs –our primary focus– but allows gauging of the economic significance of these differences.

mostly white tracts significant. Comparing Figure 3 (upper panel) and Figure 4 show that large price increases closely coincide with large white inflows in the same mostly white tracts with high predicted increase in approval rate. In the rest of tracts, they are no statistically significant differences in price increases between metro areas with high and low predicted increase in approval rates. However, in metropolitan areas with a high predicted increase in approval rates, prices increased significantly more in tracts with a small fraction of Blacks than in tracts with a large fraction of Blacks: prices increased 53% in tracts with 2.5% of Blacks, and prices increased only 38% in tracts with 95.4% of Blacks, a 15 percentage point difference.

In mostly white tracts, figure A3 of the appendix suggests a relatively higher price increase and also relatively more housing unit construction in metropolitan areas with high predicted approval rate increase. The number of housing units grew 10 and 20 percentage points more in tracts with less than 10% of blacks in such metropolitan areas compared to metropolitan areas with low predicted approval rate increase. This graph, combined with the evidence on prices, suggests that, although housing supply is elastic, the elasticity of housing supply in mostly white areas is sufficiently low to generate the price effects illustrated in figure 4 and thus the segregative effects of credit supply.

A Model of Neighborhood Choice with Borrowing Constraints

The appendix formalizes the general equilibrium mechanism described previously, in a model of neighborhood choice with credit constraints in a city with black and white populations and two neighborhoods. The two neighborhoods differ by their initial fraction of whites. This specification encompasses (a) the case of two neighborhoods with one mostly white (MW) and the other one racially mixed (Mixed), as well as (b) the case of two neighborhoods with one racially mixed (Mixed) and one mostly black neighborhood (MB).[43]

For a two-neighborhood city, Proposition A1 establishes that a relaxation of lending standards leads to a decline in black demand for the neighborhood with a larger fraction of white households if the elasticity of black demand for that neighborhood with respect to lending

[43] The model written in appendix deals with only two neighborhoods but is written under very general assumption. While tractable, the model leads to results that are both general and intuitive. An alternative could be to consider three neighborhoods – mostly whites, mostly black, and mixed – but keeping such model tractable would require very special assumption, leading to less general and robust results.

standards conditional on prices (*the partial equilibrium effect)* is lower than the price elasticity of black demand for the desirable neighborhood, multiplied by a weighted difference of the price elasticity with respect to lending standards across the two neighborhoods (*the general equilibrium effect*).

Applied to the case (a) of two neighborhoods: a mostly white neighborhood and a racially mixed neighborhood, and separately to the case (b) of two neighborhoods: a racially mixed neighborhood and a mostly black neighborhood, the proposition gives a rationale for our tract-level regressions findings. Applied to case (a), proposition A1 suggests that the general equilibrium effect on prices offset the partial equilibrium effect: Whites flowed into mostly white neighborhoods and, during the same period, house prices increased in those neighborhoods that experienced white inflows, pricing out black households from these mostly white neighborhoods (case (a)). The model therefore formalizes the general equilibrium interpretation of the metropolitan IV results in terms of price increases in neighborhoods where white households took advantage of the more lenient lending standards. Applied to case (b), proposition A1 explains the results that Blacks moved in to racially mixed neighborhoods in part because there was little general equilibrium impact of greater black demand on house prices in these racially mixed neighborhoods.[44]

[44] In the data, the lack of upward pressure on prices in mixed neighborhoods is likely coming from the observed white outflows from these neighborhoods.

6. Conclusion

Ambitious securitization goals for Fannie Mae and Freddie Mac as well as favorable capital treatment for GSEs, holders of GSE-issued mortgage-backed securities, and sponsors of Asset Backed Commercial Paper conduits have all contributed to the precipitous rise in securitization and the significant decline in lending standards. These consequences are disproportionately strong in metropolitan areas where financial intermediaries were liquidity constrained in the early 1990s, which gives us an instrument to test for the causal impact of lending standards (as measured by metropolitan median loan-to-income ratios and approval rates) on urban segregation (as measured by isolation and exposure of racial groups to each other). Although the downward trend in racial segregation continued during the first decade of the 21st century, our instrumental variables estimates reveal that this decline has been noticeably slower in metro areas where lending standards were relaxed the most.

The interpretation of our metro-level results as reflecting the causal effect of a supply shock in mortgage credit market on racial segregation, is reinforced by the fact that our instruments appear uncorrelated with potential demand shifters as well as with other structural determinants of racial segregation such as income inequality or crime.

The census-tract level results provide an understanding of the demographic and prices changes that underlie our metro-level results. In the areas where credit standards were more relaxed, we observe significant outflows of Whites from racially mixed to mostly white neighborhoods, whose prices increase significantly more than in other neighborhoods. Credit standards did allow Blacks to move from mostly blacks to racially mixed neighborhoods but the resulting effect of their exposure to Whites was dampened by the mobility of Whites towards mostly white neighborhoods.

Research has shown that segregation has negative effects on education (Card and Rothstein 2007; Hanushek, Kain, and Rivkin 2009) and crime (Weiner, Lutz, and Ludwig 2009). The results of this paper have noteworthy implications for any type of policy designed to foster cheaper access to credit as a means of increasing the welfare of poor and minority families. Rajan (2010) discusses how the political response to increasing income inequality led to such credit supply policies, which had the unintended consequence of unleashing the credit boom that played a major role in the financial crisis of 2008–2009. Our findings underscore another set of unintended consequences that became manifest even before the crisis: while increasing homeownership for disadvantaged groups, the relaxation of credit standards significantly aggravated racial segregation.

In that spirit, Levine, Levkov, and Rubinstein (2008) is somewhat close to our paper in that it looked at the impact bank deregulations on the labor market opportunities of black workers. Levine et al. (2008) found that credit market liberalization led to a decline in the black-white wage gap, consistent with labor economics taste-based theories of discrimination in hiring (Becker 1957). In contrast, in urban economics, models of segregation typically predict large degrees of racial segregation across neighborhoods in competitive housing markets with perfect credit markets (Benabou 1996). Hence the overall welfare effect of increased credit supply is open to debate.

Our paper uses the more recent expansion of credit in the 2000-2006, as a large scale experiment to measure the impact of the relaxation of lending standards on segregation. The empirical analysis on this period is facilitated by the existence of exhaustive data on individual mortgage applications and originations, through the Home Mortgage Disclosure Act (HMDA). Yet, other significant episodes of mortgage credit expansion in the 1920s and in the post-war period occurred during a period of increasing racial segregation. In the post-war period, white flight from central cities contributed to a significant increase in racial segregation (Boustan 2010). While it is beyond the scope of this paper, establishing the role played by mortgage credit in white flight is an interesting avenue for future research.

References

Acharya, VV, Richardson, M, and Van Nieuwerburgh, S 2011, 'Guaranteed to Fail: Fannie Mae, Freddie Mac, and the Debacle of Mortgage Finance'. Princeton University Press

Adelino, M, Schoar, A, and Severino, F, 2012, 'Credit Supply and House Prices: Evidence from Mortgage Market Segmentation,' *NBER Working Paper* 17832.

Acharya, VV, Schnabl, P, and Suarez, G 2010, 'Securitization without Risk Transfer', *NBER Working Paper 15730*.

Adrian, T and Shin, H 2010, 'The Changing Nature of Financial Intermediation and the Financial Crisis of 2007-09', *Annual Review of Economics*. pp. 603–18.

Baum-Snow, N 2007, 'Did Highways Cause Suburbanization?', *The Quarterly Journal of Economics*, 122(2) pp. 775-805.

Bayer, P, Fang, H and McMillan, R 2005, 'Separate When Equal? Racial Inequality and

Residential Segregation', *NBER Working Paper No. w11507*.

Bayer, P, McMillan, R, and Rueben, KS 2004, 'What Drives Racial Segregation? New Evidence Using Census Microdata', *Journal of Urban Economics*, vol. 56, pp. 514–535.

Becker, G 1957, *The Economics of Discrimination*, University of Chicago Press: Chicago, IL.

Benabou, R 1996, 'Equity and Efficiency in Human Capital Investment: The Local Connection', *The Review of Economic Studies*, vol. 63, no. 2, pp. 237–264.

Boustan, LP 2010, 'School Desegregation and Urban Change: Evidence from City Boundaries', *NBER Working Paper*, vol. 16434.

Brock, William A., and Steven N. Durlauf, 2001, 'Discrete choice with social interactions', *The Review of Economic Studies*, vol. 68, no. 2, pp. 235–260.

Brooks, Richard RW. 2011, 'Covenants without courts: enforcing residential segregation with legally unenforceable agreements', *American Economic Review*, vol. 101, no. 3, pp 360–265.

Bucks, Brian, and Karen M. Pence 2006, *Do homeowners know their house values and mortgage terms?* Divisions of Research & Statistics and Monetary Affairs, Federal Reserve Board.

Campbell, Sean D., et al. 2009 'What moves housing markets: A variance decomposition of the rent–price ratio', *Journal of Urban Economics*, vol. 66, no. 2, pp 90–102.

Card, D 2009, 'Immigration and Inequality', *The American Economic Review*, vol. 99, no. 2, pp. 1–21.

Card, D and Rothstein, J, 2007, 'Racial Segregation and the Black-White Test Score Gap', *Journal of Public Economics*, vol. 91, pp. 2158–2184.

Card, D, Mas, A, and Rothstein, J 2008, 'Tipping and the Dynamics of Segregation', *The Quarterly Journal of Economics*, vol. 123, no. 1, Oxford University Press, pp. 177–218.

Clotfelter, CT 1999, 'Public School Segregation in Metropolitan Areas', *Land Economics*, vol. 75, no. 4, pp. 487–504.

Cutler, David M., and Edward L. Glaeser, 1997, 'Are ghettos good or bad?', *The Quarterly Journal of Economics,* vol. 112, no. 3, pp 827–872.

Cutler, DM, Glaeser, EL, and Vigdor, JL 1999, 'The Rise and Decline of the American Ghetto', *Journal of Political Economy*, vol. 107, no. 3, pp. 1–52.

Cutler, David M., Edward L. Glaeser, and Jacob L. Vigdor 2008 'When are ghettos bad? Lessons from immigrant segregation in the United States', *Journal of Urban Economics*, vol. 63, no. 3, pp 759–774.

Dell'Arriccia, G, Igan, D, and Laeven, L 2009, 'Credit booms and lending standards: Evidence from the subprime mortgage market', *European Banking Center Working Paper*, pp. 1–44.

Department of Housing and Urban Development 1998, *U.S. Housing Market Conditions*.

DiSalvo, JV 1999, 'Federal reserve geographic banking market definitions', *Federal Reserve Bank of Philadelphia*.

Epple, D and Romano, R 2011, 'Peer effects in education: A survey of the theory and evidence', *Handbook of Social Economics*.

Epple, D and Sieg, H 1999, 'Estimating Equilibrium Models of Local Jurisdictions', *The Journal of Political Economy*, vol. 107, no. 4, pp. 645–681.

Farley, R, Steeh, C, Krysan, M, Jackson, T, and Reeves, K 1994, 'Stereotypes and Segregation: Neighborhoods in the Detroit Area', *American Journal of Sociology*, vol. 100, no. 3, pp. 750–780.

Ferreira, F and Gyourko, J 2011, 'Anatomy of the Beginning of the Housing Boom: U.S. Neighborhoods and Metropolitan Areas, 1993-2009', *NBER Working Paper*.

Garriga, C 2009, 'Lending standards in mortgage markets', *Economic Synopses*, Federal Reserve Bank of St. Louis.

Gelain, P, Lansing, KJ, and Mendicino, C 2012, 'House prices, credit growth, and excess volatility: Implications for monetary and macroprudential policy', *The International Journal of Central Banking*, forthcoming.

Glaeser, E and Vigdor, J 2012, 'The End of the Segregated Century: Racial Separation in America's Neighborhoods, 1890-2010', *Manhattan Institute for Policy Research: Civil Report*, vol. 66.

Glaeser, EL 2001, 'Racial Segregation in the 2000 Census: Promising News', *The Brookings Institution Working Papers*, pp. 1–16.

Goodman Jr, John L., and John B. Ittner 1992, 'The accuracy of home owners' estimates of house value', *Journal of Housing Economics*, vol. 2, no. 4, pp 39-357.

Goux, D and Maurin, E 2007, 'Close Neighbours Matter: Neighbourhood Effects on Early Performance at School', *The Economic Journal*, vol. 117, pp. 1–23.

Guerrieri, Veronica, Daniel Hartley, and Erik Hurst 2012, 'Within-City Variation in Urban Decline: The Case of Detroit', *American Economic Review*, vol. 102, no. 3, pp 120–26.

Gyourko, Joseph, Christopher Mayer, and Todd Sinai 2006, 'Superstar Cities', *NBER Working Paper*, no. 12355.

Hanushek, EA, Kain, JF, and Rivkin, SG 2009, 'New Evidence About Brown v. Board of Education: The Complex Effects of School Racial Composition on Achievement', *The*

Journal of Labor Economics, 27(3), pp. 349-383.

Himmelberg, Charles, Christopher Mayer, and Todd Sinai 2005, ‘Assessing High House Prices: Bubbles, Fundamentals and Misperceptions’, *Journal of Economic Perspectives*, pp 67-92.

Hoxby, CM 2001, ‘Peer Effects in the Classroom: Learning from Gender and Race Variation’, *NBER Working Paper Series*, pp. 1–64.

Favara, Giovanni, and Jean Imbs, 2010, ‘Credit Supply and the Price of Housing’, CEPR Discussion Papers 8129.

Jayaratne, Jith, and Philip E. Strahan, 1996, ‘The finance-growth nexus: Evidence from bank branch deregulation’, *The Quarterly Journal of Economics,* vol. 111, no. 3, pp 639–670.

Keys, BJ, Mukherjee, T, Seru, A, and Vig, V 2010, ‘Did Securitization Lead to Lax Screening? Evidence from Subprime Loans’, *Quarterly Journal of Economics*, vol. 125, pp. 307–362.

Krysan, Maria, and Reynolds Farley 2002, ‘The residential preferences of blacks: Do they explain persistent segregation?’ *Social Forces*, vol. 80, no. 3, pp 937–980.

Levine, R, Levkov, A, and Rubinstein, Y 2008, *Racial discrimination and competition*, NBER Working Paper No. 14273.

Loutskina, E 2011, ‘The role of securitization in bank liquidity and funding management’, *Journal of Financial Economics*, vol. 100, no. 3, pp. 663–684.

Loutskina, E and Strahan, PE 2009, ‘Securitization and The Declining Impact of Bank Finance on Loan Supply: Evidence from Mortgage Acceptance Rates’, *Journal of Finance,* vol 64, pp. 861-889.

Lutz, Byron 2011, ‘The end of court-ordered desegregation’ *American Economic Journal: Economic Policy,* vol. 3, no. 2, pp 130–168.

Manski, CF 1993, ‘Identification of Endogenous Social Effects: The Reflection Problem’, *The Review of Economic Studies*, vol. 60, no. 3, pp. 531–542.

Massey, Douglas S 2004, ‘Segregation and stratification’, *Du Bois Rev,* vol. 1, no. 1, pp 7–25.

Massey, DS and Denton, NA 1988, ‘The Dimensions of Residential Segregation’, *Social Forces*, vol. 67, no. 2, pp. 281–315.

Massey, Douglas S., Gretchen A. Condran, and Nancy A. Denton. "The effect of residential segregation on black social and economic well-being." *Social Forces* 66.1 (1987): 29-56.

Massey, Douglas S., Michael J. White, and Voon-Chin Phua, 1996, ‘The dimensions of segregation revisited’, *Sociological Methods & Research*, vol. 25, no. 2, pp 172–206.

Mian, A and Sufi, A 2009, ‘The Consequences of Mortgage Credit Expansion: Evidence from the U.S. Mortgage Default Crisis’, *Quarterly Journal of Economics*, pp. 1–49.

Mian, A, Sufi, A, and Trebbi, F 2011, 'Foreclosures, House Prices, and the Real Economy', *NBER Working Paper*.

Munnell, AH, Tootell, GMB, Browne, LE and McEneaney, J 1996, 'Mortgage Lending in Boston: Interpreting HMDA Data', *The American Economic Review*, vol. 86, no. 1, pp. 25–53.

Ng, J, 2012, 'Did Securitization Cause the Credit Boom', mimeo Northwestern University

Office of Federal Housing Enterprise Oversight 1998, *Unified Agenda of Federal Regulatory and Deregulatory Actions*.

Ouazad, A and Rancière, R, 2013, 'General Equilibrium Effects of Credit Markets on Urban Segregation', *mimeo INSEAD*.

Purnanandam, A, 2011, 'Originate-to-Distribute Model and the Sub-prime Mortgage Crisis' *Review of Financial Studies*, vol 24, pp. 1881-1915

Rajan, RG, 2010, 'Fault Lines: How Hidden Fractures Still Threaten the World Economy', *Princeton University Press*.

Ramcharan, R, Van den Heuvel, S, and Verani, S 2012, 'From Wall Street to Main Street: The Impact of the Financial Crisis on Consumer Credit Supply', *Federal Reserve Board Working Paper*.

Raphael, Steven, 1998, 'The spatial mismatch hypothesis and black youth joblessness: evidence from the San Francisco Bay Area', *Journal of Urban Economics,* vol. 43, no. 1, pp 79–111.

Ross, SL and Yinger, J 2002, *The Color of Credit: Mortgage Discrimination, Research Methodology, and Fair-Lending Enforcement*, The MIT Press.

Glaeser, Edward L., Joseph Gyourko, and Albert Saiz, 2008 'Housing supply and housing bubbles', *Journal of Urban* Economics, vol. 64, no. 2, pp 198–217.

Saiz, A, 2010, 'The Geographic Determinants of Housing Supply', *The Quarterly Journal of Economics*, pp. 1253–1296.

Schelling, Thomas C., 1971 'Dynamic models of segregation', *Journal of mathematical sociology,* vol. 1, no. 2, pp 143-186.

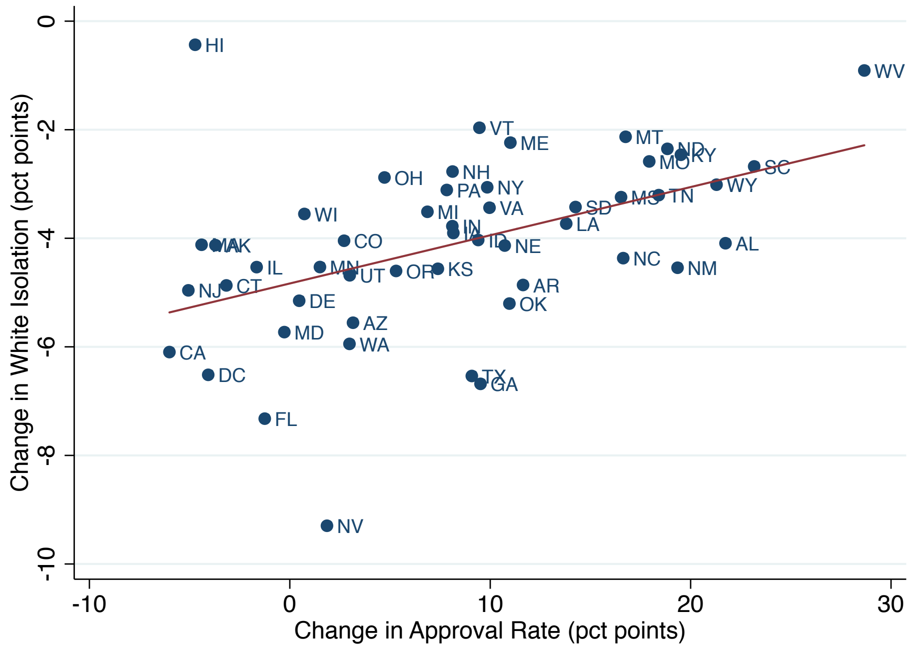
U.S. Treasury, 1990, *Report of the Secretary of the Treasury on Government Sponsored Enterprises*.

Weiner, DA, Lutz, BF, and Ludwig, J, 2009, 'The Effects of School Desegregation on Crime', *NBER Working Paper Series*.

Figure 1. Growth of Approval Rates and Isolation Changes, by State

Scatter plots of the change in isolation from 2000 to 2010 on the vertical axis, and changes in approval rates from 2000 to 2006. Fitted line is the OLS regression of isolation changes on approval rate changes. Each point on this scatter plot is a state, identified by its two-letter code. States with high Hispanic inflows: the fraction of Hispanics grew by more than the median increase across states (+2.66 percentage points).

(i) White Isolation Changes



(ii) Black Isolation Changes, by State-Level Hispanic Inflow

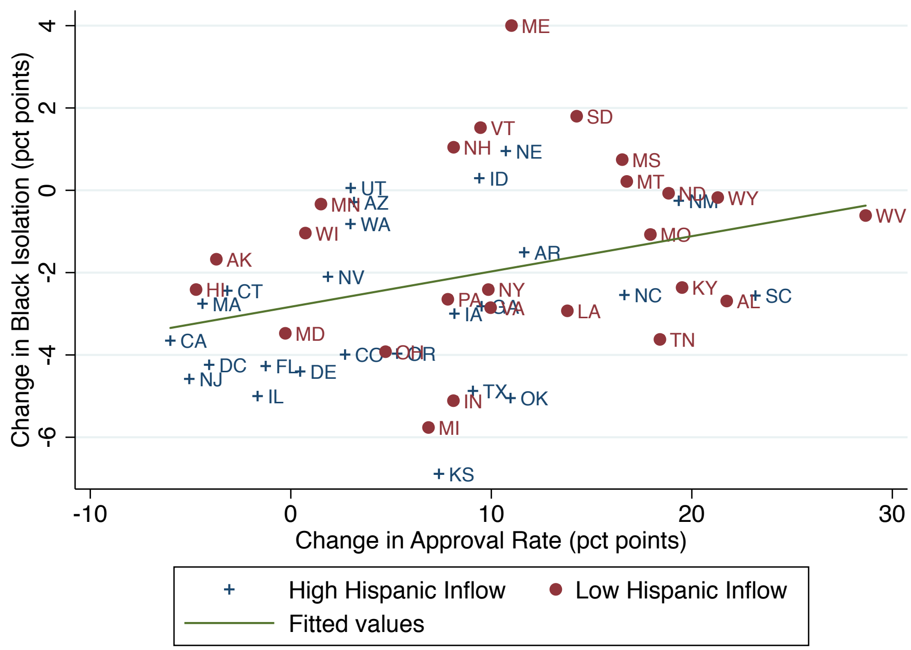
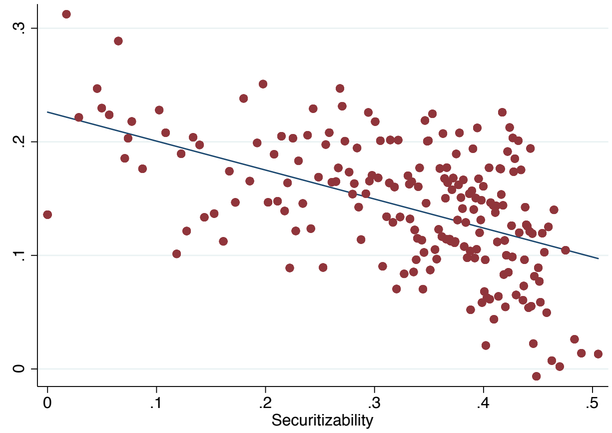


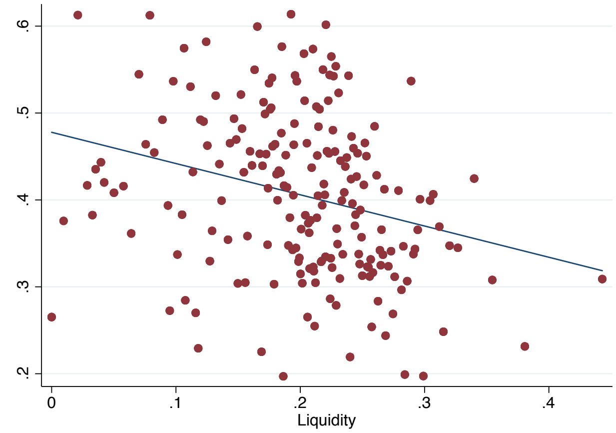
Figure 2. Lending Standards, Liquidity Measures and Securitizations

These two graphs show the negative correlation between securitizability and the growth of approval rates (in percentage points), and between liquidity and the growth of LTIs from 2000 to 2006. Each point is a metropolitan area. The blue line is the fitted line from an ordinary least squares regression with no additional covariates.



ΔApproval = 0.225 – 0.254 Securitizability + ε
R^2= 0.08, F=13.33

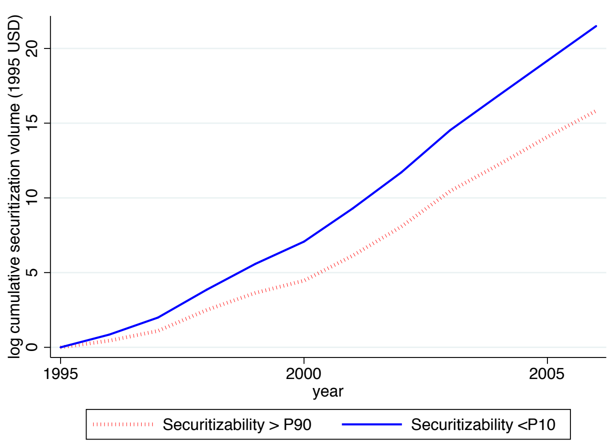
(i) Growth of approval rate 2000-2006
and Securitizability in 1990



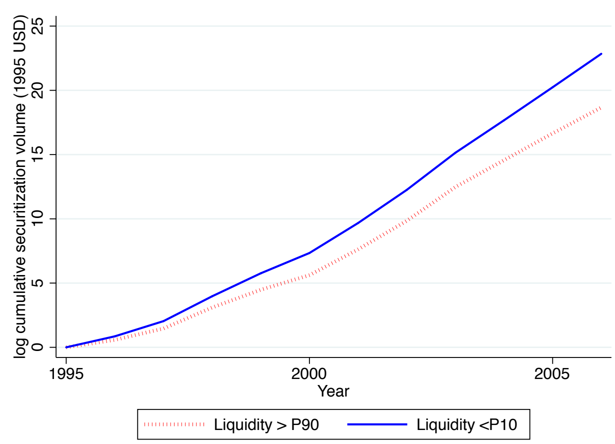
ΔLTI = 0.477 – 0.358 Liquidity + ε
R^2= 0.02, F=9.06

(ii) Loan to income ratio change 2000-2006
and Liquidity in 1990

This graph presents the log increase in securitization volume from 1995 (0%) to 2006, in constant 1995 dollars (deflated using the CPI index), for metropolitan areas with securitizability (left) and liquidity (right) below the 10^{th} percentile of the metropolitan area distribution of the liquidity measure, above the 90^{th} percentile..



(iii) Securitizations: annual log volume
by 1990 Securitizability Index



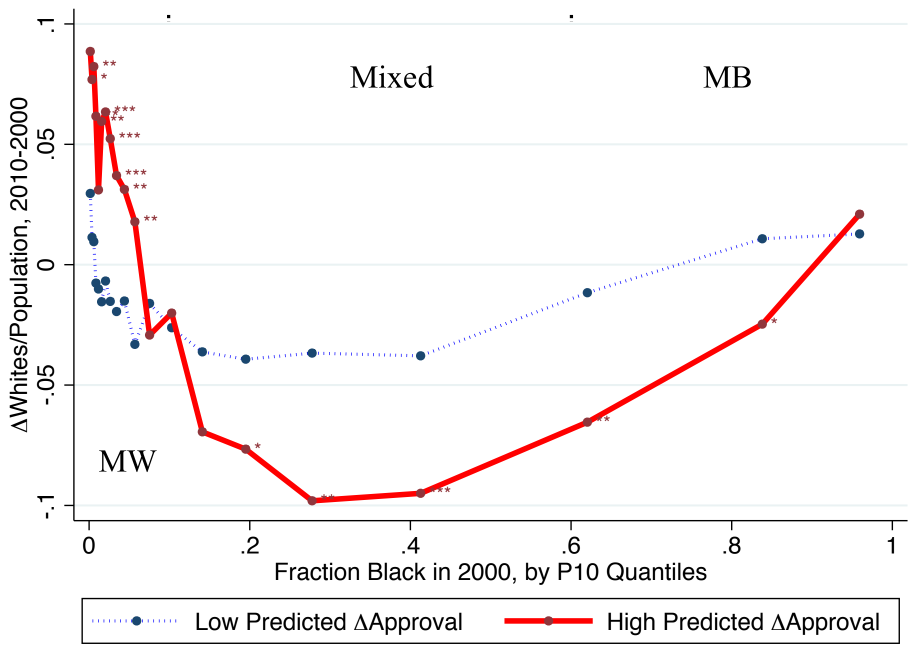
(iv) Securitizations: annual log volume,
by 1990 Liquidity Level

Figure 3. Inflows into Census Tracts, by Percentage Black in Tract

Stars next to a point indicate the significance of the F test that the coefficient for high liquidity metropolitan areas and the coefficient for low liquidity metropolitan areas are identical. One star for 10% significance, two stars for 5% significance, three stars for 1 percent. Standard errors clustered by state.

"MW": Mostly white tract, fraction black below 10%. "Mixed": Mixed tract, fraction black between 10 and 60%. "MB": Mostly black tract, fraction black above 60%.

(i) 2000-2010 Tract-Level White Population Change, by Black Fraction in Tract in 2000



(ii) 2000-2010 Tract-Level Black Population Change, by Black Fraction in Tract in 2000

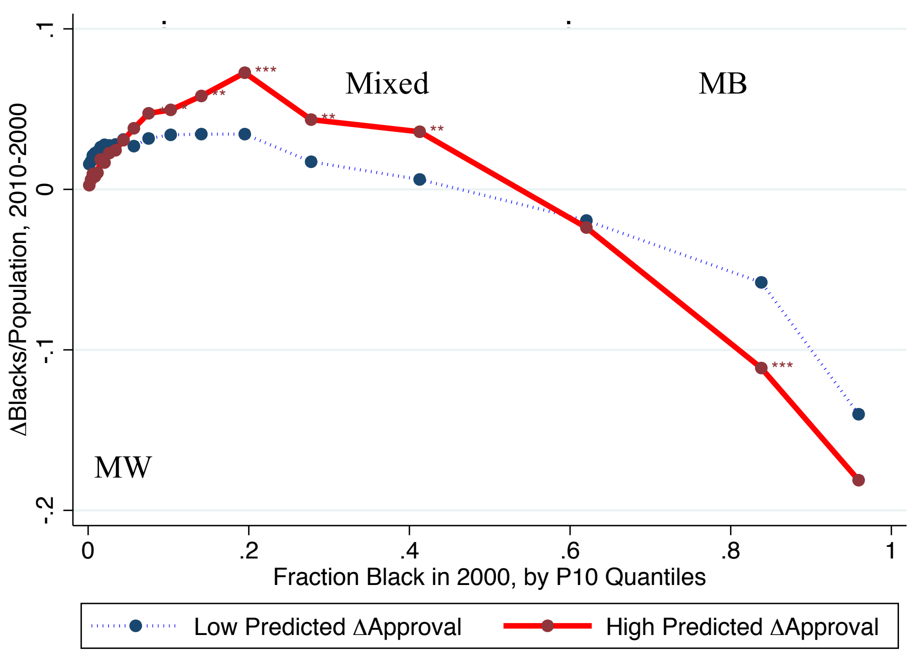


Figure 4. Price Appreciation 2000-2008 and Fraction Black in 2000

This figure plots the log house price appreciation against the fraction of black residents in 2000. House values in 2000 are taken from the Summary File 3 of the 2000 Census. House values in 2006-2010 are taken from the 2010 tract-level American Community Survey. In the American Community Survey, the median household was surveyed in 2008. Stars indicate the significance of the difference between the two curves.

****: Significant at 1%, **: Significant at 5%, *: Significant at 10%.*

High predicted ΔApproval: metropolitan level approval rates for mortgage applications increased by more than 13 percentage points. Standard errors clustered by state.

"MW": Mostly white tract, fraction black below 10%. "Mixed": Mixed tract, fraction black between 10 and 60%. "MB": Mostly black tract, fraction black above 60%.

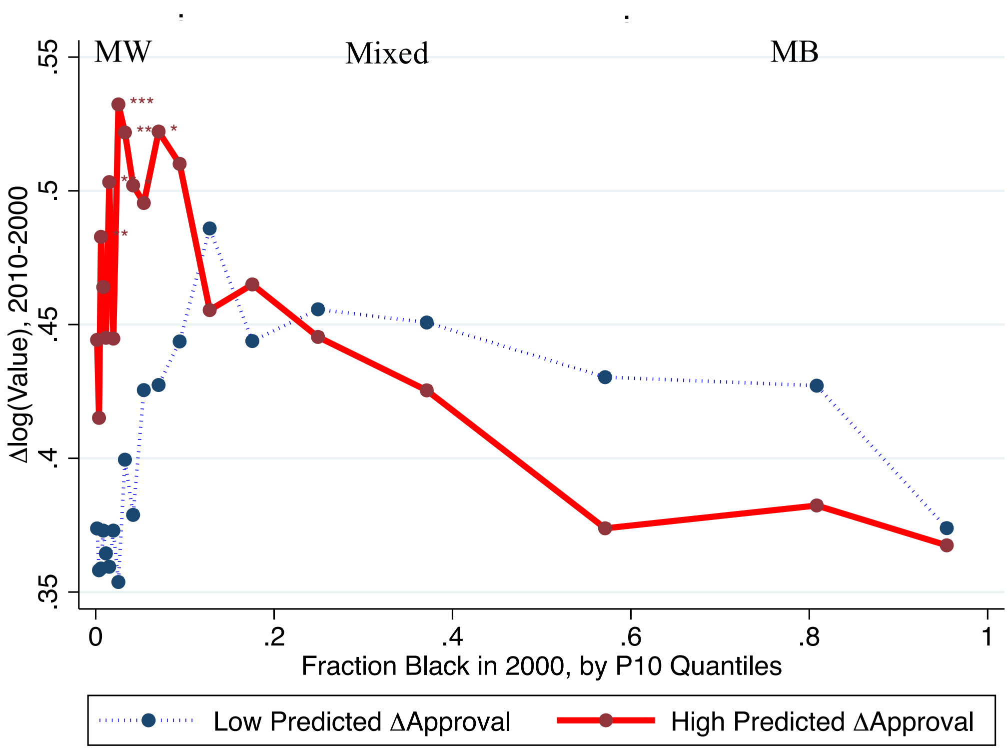


Table 1. OLS Regression of Changes in Segregation on Changes in Approval Rates and LTI

This table presents regressions of changes in black segregation (isolation and exposure) on changes in credit conditions. Regressions include state fixed effects, and standard errors are clustered at the state level. Additional controls: Elasticity of housing supply from Saiz (2010), proxies for credit risk (fraction of loans past due in 2006, fraction of high risk loans in 2006 as predicted by 1995 HMDA denial standards, median mortgage spread over treasuries in 2000 and 2006, growth in missing income loan originations). Columns (3) to (5) are estimated on Metropolitan Statistical Areas only, where (personal) income changes are from the American Community Survey and house price changes from Case-Shiller index. Demographic controls: 2000-2010 metropolitan change in fractions of Hispanic, black, Asian, and other-race populations. E. = Exposure. Case-Shiller price data are available for 206 metropolitan areas (column 3-5). Exposure and isolation results are robust to the inclusion of all metropolitan areas in the sample.

	(1) Δ Black Isolation	(2) Δ Black Isolation	(3) Δ Black Isolation	(4) Δ Black E. to Whites	(5) Δ Black E. to Hispanics
Δ Approval Rate	11.896**	11.582**	12.469**	–13.635**	0.595
	(3.423)	(2.971)	(4.255)	(3.122)	(0.766)
R-squared	0.655	0.676	0.721	0.701	0.892
Δ LTI	3.289*	3.717**	9.100*	–3.670**	–0.476
	(1.471)	(1.025)	(4.397)	(1.256)	(0.535)
R-squared	0.619	0.657	0.721	0.673	0.892
	(1) Δ Hispanic Isolation	(2) Δ Hispanic Isolation	(3) Δ Hispanic Isolation	(4) Δ Hispanic E. to whites	(5) Δ Hispanic E. to blacks
Δ Approval Rate	9.522**	8.149**	5.713*	–7.451*	–2.871**
	(3.416)	(2.897)	(2.770)	(3.635)	(0.702)
R-squared	0.772	0.797	0.850	0.798	0.749
Δ LTI	0.703	0.975	-0.034	-1.358	-0.100
	(1.005)	(1.331)	(3.348)	(1.249)	(0.554)
R-squared	0.744	0.782	0.845	0.790	0.740
	(1) Δ White Isolation	(2) Δ White Isolation	(3) Δ White Isolation	(4) Δ White E. to hispanics	(5) Δ White E. to blacks
Δ Approval Rate	2.807**	2.583**	2.423+	1.010*	–3.789**
	(0.883)	(0.870)	(1.318)	(0.421)	(0.802)
R-squared	0.931	0.933	0.958	0.948	0.809
Δ LTI	0.095	0.215	0.968	0.554**	–0.768*
	(0.419)	(0.383)	(0.902)	(0.164)	(0.317)
R-squared	0.926	0.929	0.957	0.948	0.773
Observations	939	939	939	939	939
Demographic controls	yes	yes	yes	yes	yes
State effect	yes	yes	yes	yes	yes
Additional controls	no	yes	yes	yes	yes
Income and price controls	no	no	yes	no	no

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table 2. First Stage – Growth of Approval Rates and Loan Portfolio Securitizability

This table presents the regression of changes in approval rates (in percentage points) on the loan portfolio securitizability measure (measured at the metropolitan level), for columns (1)-(4). Column (5) presents a regression of the log increase in securitization volume on securitizability. Column (3) regresses on securitizability computed using only banks whose originations in the metropolitan area represents less than 25% of their overall originations in 1990. Column (4) regresses on securitizability computed using only non-mortgage loans.

	(1)	(2)	(3)	(4)	(5)
VARIABLES	ΔApproval	ΔApproval	ΔApproval	ΔApproval	Δlog(Sec.)
Securitizability	–0.225**	–0.155**			–0.287**
	(0.062)	(0.028)			(0.091)
Securitizability Index (<25% in metro area)			–0.169**		
			(0.027)		
Securitizability Index (w.o. mortgages)				–0.352**	
				(0.067)	
Observations	939	939	939	939	939
R-squared	0.175	0.625	0.629	0.627	0.483
Demographic controls	yes	yes	yes	yes	yes
Cluster	State level	State level	State level	State level	State level
State effect	no	yes	yes	yes	yes
Additional controls	no	yes	yes	yes	yes
F stat for H_0: Securitizability Coeff. = 0	13.07	29.81	38.31	28.08	10.03

Table 3. First Stage – Growth in Loan to Income Ratio and Bank Liquidity

This table presents the regression of changes in LTI on the liquidity measure (measured at the metropolitan area level), for columns (1)-(4). Column (5) presents a regression of the log increase in securitization volume on liquidity. Column (3) regresses on liquidity computed using only banks whose originations in the metropolitan area represent less than 25% of their overall originations in 1990. Column (4) regresses on liquidity computed as (cash+securities)/(total assets excluding mortgages).

	(1)	(2)	(3)	(4)	(5)
VARIABLES	ΔLTI	ΔLTI	ΔLTI	ΔLTI	Δlog(Sec.)
Liquidity	–0.329**	–0.196**			–0.509**
	(0.113)	(0.069)			(0.174)
Liquidity (< 25% in metro area)			–0.188*		
			(0.079)		
Liquidity (w.o. mortgages)				–0.118*	
				(0.056)	
Observations	939	939	939	939	939
R-squared	0.043	0.586	0.585	0.585	0.486
Demographic controls	yes	yes	yes	yes	yes
Cluster	State level	State level	State level	State level	State level
State effect	no	yes	yes	yes	yes
Additional controls	no	yes	yes	yes	yes
F stat for H_0: Liquidity Coeff. = 0	8.50	8.07	5.63	4.53	8.57

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** $p<0.01$, * $p<0.05$, + $p<0.1$*

Table 4. IV Estimates of the Impact of Credit Supply on Segregation

Columns 1—3 of this table presents instrumental variable regressions of changes in black isolation and exposure on changes in approval rates and loan-to-income ratio, instrumented respectively by loan portfolio securitizability and by liquidity. Column 4 presents the reduced form linear regression results. Demographic controls: 2000-2010 metropolitan change in fractions of Hispanic, black, Asian, and other-race populations.

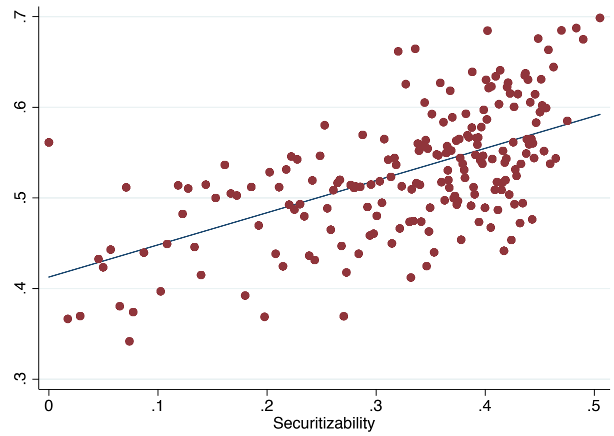
	Instrumental variable regression				Reduced form
	(1)	(2)	(3)		(4)
Dependent variable:	ΔBlack Isolation	ΔBlack E. to Whites	ΔBlack E. to Hispanics		ΔBlack Isolation
ΔApproval Rate	18.098**	–17.760**	–0.821	Securitizability	–4.798**
	(6.004)	(6.035)	(2.609)		(1.070)
F statistic	6.281	11.79	108.6	F statistic	15.67
Cragg Donald	63.94	63.94	63.94	R squared	0.16
First stage F statistic	13.07	13.07	13.07		
ΔLTI	20.249**	–12.940*	–7.047*	Liquidity	–6.142**
	(7.235)	(6.592)	(3.213)		(0.990)
F statistic	5.124	4.054	60.06	F statistic	23.48
Cragg Donald	15.25	15.25	15.25	R squared	0.15
First stage F statistic	8.501	8.501	8.501		
Dependent variable:	Δ Hispanic Isolation	Δ Hispanic E. to whites	Δ Hispanic E. to blacks		ΔHispanic Isolation
ΔApproval Rate	–4.288	2.204	1.834	Securitizability	1.244+
	(3.020)	(3.599)	(1.968)		(0.691)
F statistic	60.61	42.58	96.08	F statistic	62.89
Cragg Donald	63.94	63.94	63.94	R squared	0.65
First stage F statistic	13.07	13.07	13.07		
ΔLTI	–7.962+	10.816+	–1.410	Liquidity	2.629*
	(4.077)	(5.939)	(2.664)		(1.175)
F statistic	21.28	48.83	75.09	F statistic	55.52
Cragg Donald	15.25	15.25	15.25	R squared	0.65
First stage F statistic	8.501	8.501	8.501		
Dependent variable:	ΔWhite Isolation	ΔWhite E. to Hisp.	ΔWhite E. to Blacks		ΔWhite Isolation
ΔApproval Rate	–0.889	2.922**	–1.026+	Securitizability	0.168
	(1.464)	(0.998)	(0.614)		(0.427)
F statistic	256.1	763.9	160.0	F statistic	226
Cragg Donald	63.94	63.94	63.94	R squared	0.80
First stage F statistic	13.07	13.07	13.07		
ΔLTI	0.188	2.270*	–0.775	Liquidity	–0.083
	(1.777)	(0.990)	(1.040)		(0.623)
F statistic	234.8	103.8	175.2	F statistic	224.71
Cragg Donald	15.25	15.25	15.25	R squared	0.80
First stage F statistic	8.501	8.501	8.501		
Demographic controls	yes	yes	yes	Demographic controls	yes
Observations	939	939	939	Observations	939

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

ONLINE APPENDIX

Figure A1. Level of Approval Rates and Securitizability

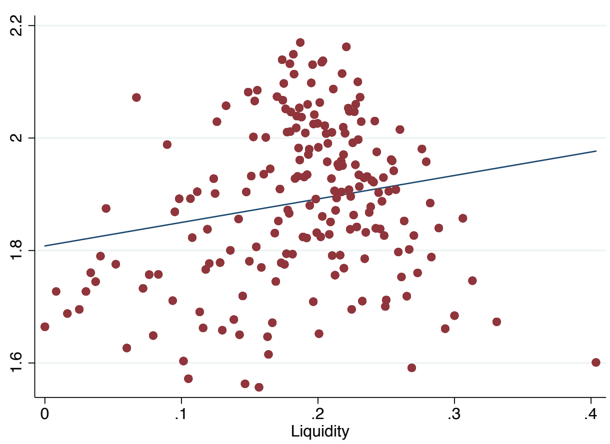
The graph shows the positive relationship between loan portfolio securitizability in 1990-1994 and the level of approval rates in 2000. Each point is a metropolitan area (MSA or μSA). The blue line is the fitted line from an ordinary least squares regression with no additional covariates.



Approval = 0.413 + 0.353 Securitizability
R^2= 0.10, F=20.17

Figure A2. Level and of Loan-to-income Ratios and Liquidity

The graph shows the positive relationship between bank liquidity in 1990-1994 and loan-to-income ratios in 2000. Each point is a metropolitan area (MSA or μSA). The blue line is the fitted line from an ordinary least squares regression with no additional covariates.



LTI = 1.866 + 0.102 Liquidity + ε
R^2= 0.01, F=0.28

Figure A3. Tract-Level Construction and Metropolitan Lending Standards

*The graph (described in section 5.1) shows the tract-level percentage change in housing units as a function of the fraction of black residents in 2000. Each point is a coefficient in the regression of population change on the dummy for each bin. Tracts are sorted in 20 bins of equal size, sorted by increasing fraction of blacks in the tract in 2000. Stars indicate the significance of the F test that the coefficient of the red curve (high predicted approval rate increase) is equal to the coefficient of the blue curve (low predicted approval rate increase) for the same fraction of black residents. *** : significant at 1%, ** : significant at 5%, * : significant at 10%.*

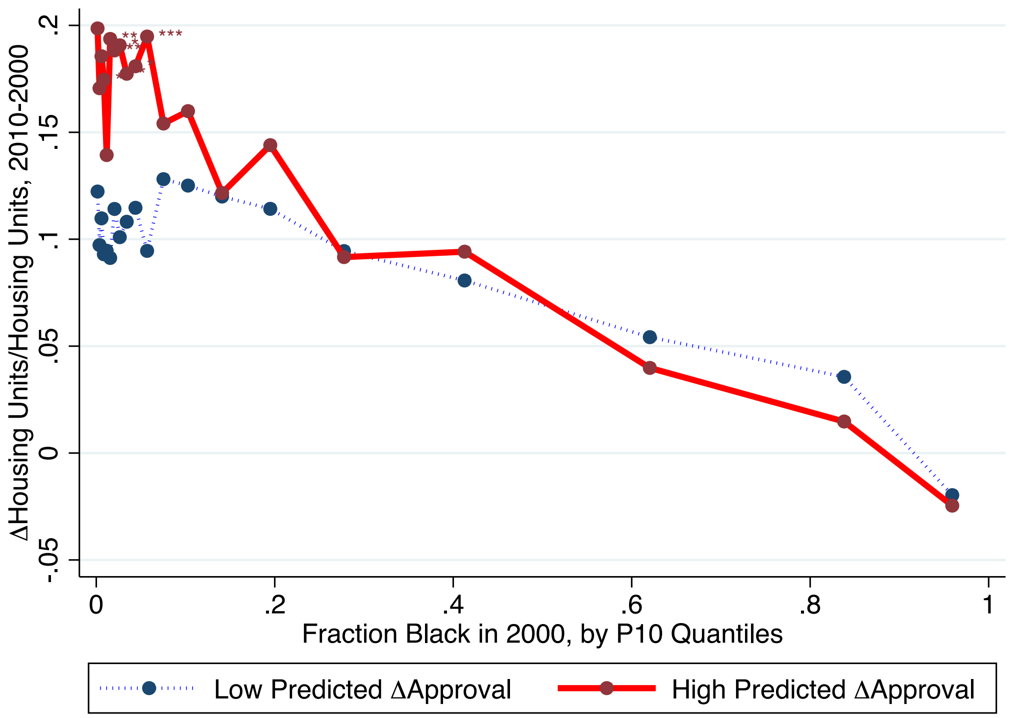


Table A1. Descriptive Statistics – Demographics

This table presents population and black segregation statistics for the 366 Metropolitan Statistical Areas. Isolation and exposure are computed across census tracts, following the definitions of Cutler, Glaeser, and Vigdor (1999). The data is from the Summary File I of the 2000 and 2010 Census, with consistent 2000 census tract borders. ppt: percentage points.

	2000		2010		2000-2010 Change		# of MSAs with	
	Average	Median	Average	Median	Average	Median	increases (%)	declines (%)
MSA population	636,845	217,874	705,695	250,093	+68,850	+21,503	324 88.5%	42 11.5%
Black Population in MSA	81,582	15,734	91,537	17,923	9,955	1,900	326 89.1%	40 10.9%
Fraction Black in MSA	12.3%	6.4%	12.5%	7%	+0.2 ppt	+0.3 ppt	268 73.2%	98 26.8%
Hispanic Population in MSA	89,347	10,900	127,607	18,843	+38,259	+7,752	366 100.0%	0 0.0%
Fraction Hispanic in MSA	13.1%	3.7%	17.1%	6.2%	+4.0 ppt	+2.5 ppt	366 100.0%	0 0.0%
Black Isolation	50.5%[a]	21.0%	45.4%[a]	18.5%	–5.1 ppt	–1.5 ppt	93 25.4%	273 74.6%
Black Exposure to Hispanics	11.1%[a]	5.5%	14.8%[a]	8.4%	+3.7 ppt	+2.8 ppt	359 98.1%	7 1.9%
Black Exposure to Whites	33.2%[a]	58.5%	32.7%[a]	55.3%	–0.5 ppt	–2.0 ppt	77 21.0%	289 79.0%

[a]: such average is weighted by the metropolitan area's black population.

Table A2. Credit Conditions

This table presents average mortgage application and origination characteristics from 2000 to 2006, as well as price changes, for all metropolitan areas. Data sources: Mortgage data from HMDA 2000-2006, Case Shiller Weiss house price index 2000-2006 from Standard and Poor's, housing supply elasticity data from satellite data and the Wharton Land Use Regulation Index (Saiz 2010). *Metropolitan areas: Metropolitan Statistical Areas (MSAs) and Micropolitan Statistical Areas (μSAs).*

	2000		2006		2006-2000	
	Mean	S.d.	Mean	S.d.	Mean	S.d.
All Borrowers						
Approval Rate (Pct.)	70.13	11.56	84.2	4.63	14.07	10.42
Loan-to-income ratio (Median)	1.89	0.27	2.3	0.38	0.41	0.19
- in larger metro areas (366 MSAs)	2.03	0.26	2.46	0.36	0.44	0.18
- in metro areas with housing supply elasticity above median	2.01	0.26	2.44	0.37	0.43	0.18
Mortgage originations (M USD)	564	2331	957	3914	53.83%	33.07%
Case Shiller Index (100=1995, Q1)	122	10	192	51	42.59%	23.09%
Mortgage originations (Number)	4,437	15,129	5,680	19,646	17.06%	30.73%
White Borrowers						
Approval Rate (Pct.)	75.56	10.14	86.39	4.3	10.83	9.12
Loan-to-income ratio (Median)	1.89	0.27	2.28	0.38	0.4	0.19
Black Borrowers						
Approval Rate (Pct.)	63.62	21.42	76.84	15.96	13.08	23.42
Loan-to-income ratio (Median)	2.02	0.45	2.42	0.5	0.42	0.52
Hispanic Borrowers						
Approval Rate (Pct.)	68.06	19.79	80.65	12.39	12.74	22.79
Loan-to-income ratio (Median)	2.02	0.44	2.41	0.5	0.41	0.43
Number of metropolitan areas	940		940		940	

Table A3. Confounding Factors – Correlation of Changes in Approval Rates, Changes in the Loan-to-Income ratio, with Income Changes.

This table presents regressions of log wage or personal income changes between 2000 and 2006 and changes in the fraction of Whites, Hispanics, and Blacks (in percentage points), on changes in credit conditions over the same time period. The data is at the metropolitan area (MSA) level, where wage and personal income changes can be estimated from the 2000 and 2006 American Community Survey.

	Regressor:		
Dependent variable:	(1) ΔApproval Rate	(2) ΔLTI	Observations
Income			
Δlog(income) *Wage income*	–0.067 (0.072)	0.134** (0.031)	230
Δlog(income) *Pers. income*	–0.140+ (0.072)	0.139** (0.035)	230
Black Δlog(income) *Wage income*	–0.085 (0.241)	0.369* (0.159)	230
Hispanic Δlog(income) *Wage income*	–0.509* (0.206)	0.183+ (0.097)	230
White Δlog(income) *Wage income*	–0.129* (0.057)	0.106** (0.034)	230
Asian Δlog(income) *Wage income*	–0.384 (0.386)	0.147 (0.249)	230
Black-White *log Wage Income gap*	–0.019 (0.123)	–0.064 (0.074)	230
Demographics			
ΔPct. White 2000-2010	0.056** (0.014)	0.002 (0.011)	939
ΔPct. Hispanics 2000-2010	–0.024+ (0.014)	0.002 (0.012)	939
ΔPct. Blacks 2000-2010	–0.024** (0.007)	–0.002 (0.003)	939

Robust standard errors clustered by state in parenthesis.
*Significance level: ** p<0.01, * p<0.05, + p<0.1*

Table A4.1. Correlation of the Instruments with Potential Demand Shifters

This table presents regressions of log wage or personal income changes between 2000 and 2006 and changes in the fraction of Whites, Hispanics, and Blacks (in percentage points), on our two instruments, liquidity and loan portfolio securitizability. The data is at the metropolitan area (MSA) level, where wage and personal income changes can be estimated from the 2000 and 2006 American Community Survey.

	Regressor:		
Dependent variable:	(1) Liquidity	(2) Securitizability	Observations
Income			
Δlog(income) *Wage income*	–0.256 (0.187)	–0.033 (0.095)	230
Δlog(income) *Pers. income*	–0.319 (0.193)	0.007 (0.105)	230
Black Δlog(income) *Wage income*	0.679 (0.525)	0.141 (0.496)	230
Hispanic Δlog(income) *Wage income*	0.481 (0.593)	0.576 (0.414)	230
White Δlog(income) *Wage income*	–0.220 (0.172)	0.041 (0.096)	230
Asian Δlog(income) *Wage income*	–0.907 (0.795)	0.533 (0.565)	230
Black-White log wage income gap	0.363 (0.275)	–0.251 (0.182)	230
1990 P90/P10 Wage Income Ratio	–1.183 (8.210)	–4.947 (4.965)	230
Demographics			
ΔPct. White 2000-2010	0.004 (0.015)	–0.003 (0.010)	939
ΔPct. Hispanics 2000-2010	–0.015 (0.016)	–0.004 (0.010)	939
ΔPct. Blacks 2000-2010	0.005 (0.005)	0.005 (0.005)	939

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table A4.2. Correlation of the Instruments with Other Confounders

This table presents regressions of log wage or personal income changes between 2000 and 2006 and changes in the fraction of Whites, Hispanics, and Blacks (in percentage points), on our two instruments, liquidity and loan portfolio securitizability. The data is at the metropolitan area (MSA) level, where wage and personal income changes can be estimated from the 2000 and 2006 American Community Survey.

	Regressor: (1)	(2)		
Dependent variable:	Liquidity	Securitizability	Correlation with Black Isolation in 2000	Observations
College Premium	0.165 (0.227)	–0.419 (0.291)	0.001+ (0.005)	266
Average July Temperature	-2.803 (9.295)	2.807 (9.756)	0.285** (0.047)	919
Income inequality Gini, 2000	-0.055 (0.081)	0.126 (0.281)	0.001 (0.001)	266
Change in Gini 2000-2010	-0.192 (0.182)	0.098 (0.125)	0.001+ (0.000)	266
Burglary Rate Per 100,000 residents	–0.003 (0.003)	0.001 (0.003)	0.086*** (0.006)	919
Homicide rate Per 100,000 residents	0.001 (0.001)	0.001 (0.001)	7.411*** (0.637)	919

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table A5. Predicting Hispanic Population Growth with 1980 Metropolitan Hispanic Population

This table presents regressions of the impact of changes in credit standards on segregation changes, instrumenting Hispanic demographic changes by the fraction of Hispanics in the metropolitan area in 1980 (column (1)) and instrumenting approval rate growth (resp., LTI growth) by securitizability (resp., liquidity), or IV regressions instrumenting approval rate growth (resp., LTI growth) by securitizability (resp., liquidity) with a control for the predicted increase in the fraction of Hispanics, predicted by the fraction of Hispanics in 1980 (columns (2)-(4)). The F stats are for the second stage IV regression.

VARIABLES	(1) ΔBlack Isolation	(2) ΔBlack ΔIsolation	(3) ΔE. of Blacks to Whites	(4) ΔE. of Blacks to Hispanics
ΔApproval Rate	18.098**	16.775**	–18.366+	0.935
	(6.004)	(5.815)	(9.551)	(5.693)
Δ Frac. Hispanics	–0.047			
	(0.117)			
Predicted Δ Frac. Hispanics		0.474	–0.357	0.100
		(0.364)	(0.551)	(0.366)
IV for Approval Rate	Securitizability	Securitizability	Securitizability	Securitizability
IV for Δ Frac. Hispanics	1980 Hispanic fraction	-	-	-
Observations	939	939	939	939
F stat	6.281	12.22	2.786	6.246
Cragg Donald	63.94	63.13	63.13	63.13
Angrist-Pischke	13.07	11.38	11.38	11.38

VARIABLES	(1) ΔBlack Isolation	(2) ΔBlack Isolation	(3) ΔE. of Blacks to Whites	(4) ΔE. of Blacks to Hispanics
ΔLTI	20.249**	15.889**	–17.066*	1.608
	(7.235)	(5.723)	(7.975)	(4.594)
Δ Frac. Hispanics	–0.335			
	(0.212)			
Predicted Δ Frac. Hispanics		0.481	–0.362	0.107
		(0.631)	(0.694)	(0.371)
IV for Approval Rate	Liquidity	Liquidity	Liquidity	Liquidity
IV for Δ Frac. Hispanics	1980 Hispanic fraction	-	-	-
Observations	939	939	939	939
F stat	5.124	5.919	2.711	7.078
Cragg Donald	15.25	18.47	18.47	18.47
Angrist-Pischke	8.501	9.973	9.973	9.973

Robust standard errors clustered by state in parenthesis.
*** : significant at 1%. *: significant at 5%. +:significant at 10%.*

Table A6. The Role of Housing Supply Elasticity -- Metropolitan Level Results

This table presents the main baseline IV regressions, this time with the approval rate change (resp. the LTI ratio change) interacted with the metropolitan area housing supply elasticity. The instruments are the securitizability index (resp. the liquidity measure) interacted with housing supply elasticity. Housing supply elasticity was derived in Saiz (2010).

Dependent variable:	(1) ΔBlack Isolation	(2) ΔBlack E. to whites	(3) ΔBlack E. to hispanics	(4) ΔBlack Isolation
ΔApproval Rate	8.562*	-7.700*	-2.239	
	(3.541)	(3.887)	(2.277)	
ΔApproval Rate x Housing Supply Elasticity	-4.803**	4.655**	-0.349	
	(1.056)	(1.104)	(0.413)	
Securitizability				-3.626**
				(0.934)
Securitizability x Housing Supply Elasticity				-1.595**
				(0.242)
Observations	918	918	918	919
F stat	10.47	23.47	79.15	17.37
Cragg Donald	29.26	29.26	29.26	*n.a.*
Angrist-Pischke for Approval Rate	7.878	7.878	7.878	*n.a.*

Dependent variable:	(1) ΔBlack Isolation	(2) ΔBlack E. to whites	(3) ΔBlack E. to hispanics	(4) ΔBlack Isolation
ΔLTI	11.742*	-4.320	-8.386*	
	(5.230)	(5.538)	(3.976)	
ΔLTI x Housing Supply Elasticity	-1.421**	1.577**	-0.370	
	(0.374)	(0.418)	(0.270)	
Liquidity				-4.388**
				(0.920)
Liquidity x Housing Supply Elasticity				-2.457**
				(0.419)
Observations	918	918	918	919
F stat	10.63	18.32	44.47	21.59
Cragg Donald	4.133	4.133	4.133	*n.a.*
Angrist-Pischke for Approval Rate	2.562	2.562	2.562	*n.a.*

Table A7. Black Isolation, Credit Supply, and Foreclosure Rates

This table presents regressions of the impact of changes in credit standards on segregation changes, controlling for foreclosure sales per 100 resident. Foreclosure data is from ZIP code level RealtyTrac data aggregated at the CBSA level. All regressions include a state fixed effect. Additional controls are defined in the notes of Table 3. Demographic controls: 2000-2010 CBSA-level change in fractions of Hispanic, black, Asian, and other-race populations.

VARIABLES	(1) ΔBlack Isolation	(2) ΔExposure to Whites	(3) ΔExposure to Hispanics
ΔApproval Rate	10.935**	–12.486**	–0.397
	(3.318)	(3.457)	(0.986)
Foreclosures sales per 100 resident	–0.107	0.249	–0.207*
	(0.238)	(0.270)	(0.102)
Demographic controls	yes	yes	yes
Observations	939	939	939
R-squared	0.678	0.705	0.899
Cluster	State level	State level	State level
State effect	yes	yes	yes
Additional controls	yes	yes	yes
F	74.82	14.19	327.1

VARIABLES	(1) ΔBlack Isolation	(2) ΔExposure to Whites	(3) ΔExposure to Hispanics
ΔLTI	3.580**	–3.388**	–0.631
	(0.910)	(0.907)	(0.469)
Foreclosures sales per 100 resident	–0.489*	0.686*	–0.194*
	(0.236)	(0.282)	(0.075)
Demographic controls	yes	yes	yes
Observations	939	939	939
R-squared	0.667	0.688	0.899
Cluster	State level	State level	State level
State effect	yes	yes	yes
Additional controls	yes	yes	yes
F	43.23	10.45	378.7

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table A8. Impact of Lending Standards on Segregation Measured in 2006-2010 American Community Survey Data

This table presents results of the estimation of the impact of LTI and approval rate growth on black isolation and exposure to races, using segregation measures from the 2006-2010 tract-level American Community Survey. E. = Exposure. Reading: an increase of the median LTI by 0.1 increases black isolation by 0.41 percentage point (0.1×4.136). An increase of the approval rate by 10 percentage points increases black isolation by 1.1 percentage points (0.10×11.061).

Additional controls: metropolitan area housing supply elasticity. Income and price controls: Current Population Survey white, black, Hispanic median income and 2000-2006 log change in metropolitan Case-Shiller index (columns 3—5).

VARIABLES	(1) Δ Black Isolation	(2) Δ Black Isolation	(3) Δ Black Isolation	(4) Δ Black E. to whites	(5) Δ Black E. to hispanics
Δ Approval Rate	11.061**	11.110**	14.137**	-16.392**	2.071
	(3.981)	(3.336)	(4.720)	(5.406)	(1.476)
Observations	939	939	939	939	939
R-squared	0.579	0.605	0.719	0.746	0.858
Cluster	State level	State level	State level	State level	State level
State effect	yes	yes	yes	yes	yes
Additional controls	no	yes	yes	yes	yes
Income and price controls	no	no	yes	yes	yes
F statistic	32.07	74.59	64.99	71.55	82.09

VARIABLES	(1) Δ Black Isolation	(2) Δ Black Isolation	(3) Δ Black Isolation	(4) Δ Black E. to whites	(5) Δ Black E. to hispanics
Δ LTI	4.136**	4.300**	8.759*	-8.725*	0.555
	(1.428)	(1.179)	(3.799)	(3.495)	(1.323)
Observations	939	939	939	939	939
R-squared	0.549	0.588	0.709	0.729	0.856
Cluster	State level	State level	State level	State level	State level
State effect	yes	yes	yes	yes	yes
Additional controls	no	yes	yes	yes	yes
Income and price controls	no	no	yes	yes	yes
F statistic	21.18	34.57	89.57	40.88	151.2

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table A9. Credit Supply and School Segregation, 2000 to 2006.
Changes in Black Students' School-Level Isolation

This table presents regressions of the impact of changes in credit standards on school segregation. School data is from the Common Core of Data from the U.S. Department of Education 2000 to 2006. Isolation of black Students is the average fraction of same-race peers at school for black students, for K-12 students. Similar results obtain when focusing on K-5 or K-8 students. ΔIsolation is the change in black isolation across schools within each metropolitan area. Columns (1)-(2) and (4)-(5) are OLS regressions. Column (3) is the IV regression with the change in approval rate instrumented by loan portfolio securitizability and the change in LTI instrumented by liquidity. Columns (2) and (4)-(5) include state fixed effects and additional controls defined in Table 1. Demographic controls: 2000-2010 metropolitan area change in fractions of Hispanic, black, Asian, and other-race populations.

VARIABLES	(1) ΔBlack Isolation	(2) ΔBlack Isolation	(3) ΔBlack Isolation	(4) ΔBlack Exposure to Whites	(5) ΔBlack Exposure to Hispanics
ΔApproval Rate	5.211*	8.072*	13.876**	–10.666**	1.157
	(2.072)	(3.802)	(5.173)	(3.422)	(1.063)
Demographic controls	yes	yes	Yes	yes	yes
Observations	939	939	939	939	939
R-squared	0.469	0.710	0.309	0.732	0.906
IV	no	no	yes	no	no
Cluster	State	State	State	State	State
State effect	no	yes	no	yes	yes
Additional controls	no	yes	no	yes	yes
F statistic	19.88	97.06	9.594	73.22	63.80

VARIABLES	(1) ΔBlack Isolation	(2) ΔBlack Isolation	(3) ΔBlack Isolation	(4) ΔBlack Exposure to Whites	(5) ΔBlack Exposure to Hispanics
ΔLTI	5.839**	2.381	9.634*	–2.584	–0.034
	(1.519)	(1.890)	(4.002)	(2.174)	(0.670)
Demographic controls	yes	yes	yes	yes	yes
Observations	939	939	939	939	939
R-squared	0.535	0.698	0.234	0.714	0.906
IV	no	no	yes	no	no
Cluster	State	State	State	State	State
State effect	no	yes	no	yes	yes
Additional controls	no	yes	no	yes	yes
F	67.03	57.42	10.91	72.45	68.27

Robust standard errors clustered by state in parenthesis.
*Significance levels: ** p<0.01, * p<0.05, + p<0.1*

Table A10. Income Segregation, Racial Segregation, and Lending Standards

This table presents linear regressions of (1) the level of income segregation in 2000 on black isolation in 2000 in level, (2) the change in income segregation between 2000 and 2010 on the change in black isolation during the same time period, (3) the change in income segregation on the change in approval rate between 2000 and 2006, and (4) the change in income segregation on the change in the loan-to-income ratio between 2000 and 2006. Income segregation is the ratio of the between census tract variance of household income to the total variance of household income within each metropolitan area. The data source for income segregation measures is the summary file 3 of the 2000 and 2010 censuses.

VARIABLES	(1) Income Segregation	(2) ΔIncome Segregation	(3) ΔIncome Segregation	(4) ΔIncome Segregation
Black Isolation, 2000	0.215**			
	(0.025)			
ΔBlack Isolation		-0.001		
		(0.052)		
ΔApproval Rate			-0.875	
			(2.035)	
ΔLTI				-0.568
				(0.957)
Observations	366	366	366	366
Demographic Controls	Yes	Yes	Yes	Yes
Data Set	MSAs	MSAs	MSAs	MSAs
State Effect	yes	Yes	Yes	Yes
F statistic	45.56	3.774	3.554	3.166

Online Appendix (Continued)

Appendix: Equilibrium Model of Neighborhood Choice with Borrowing Constraints

We present here a simple model with two neighborhoods (as in Benabou 1996) that shows that (i) the impact of lending standards relaxation is theoretically ambiguous, and (ii) explains how this impact depends on households' preferences, the elasticity of housing supply, and the relative value of neighborhoods. A more elaborate model is presented in a companion paper (Ouazad and Rancière 2013).

The city is comprised of two neighborhoods j=1,2. The price of housing in neighborhood j is p_j. In each neighborhood, s_j identical houses are supplied by competitive developers, so that the supply of housing in neighborhood j is increasing with the price of housing $s_j(p_j)$. Each house is inhabited by one household, and a household resides in one house. Households can choose to live outside the city. Households can be either *r=white*, or *r=black*. Race r's demand $D_j^r(p_1,p_2,W_1,W_2;\alpha)$ for housing in neighborhood j is a decreasing function of the price of housing in neighborhood j, an increasing function of the price of housing in the other neighborhood –j. When households have a preference for neighbors of a particular race, e.g. a preference for same-race neighbors, demand D_j^r also depends on the fraction of Whites W_j in each neighborhood. Finally, demand for neighborhood j is determined by the availability of credit, which depends on lending standards, which we summarize in a scalar parameter α. In a more elaborate setup (Ouazad and Rancière 2013), the lending standard parameter α captures the sensitivity of mortgage approval rates to the loan-to-income ratio.

The discussion of Section 5 of the paper can be formalized as follows. When the lending constraint is relaxed through an increase in α, Blacks' demand for housing in neighborhood 1 changes due to both increased credit supply $d\alpha$ and due to increased prices $dp_k/d\alpha$:

$$dDj^b/d\alpha = \partial D_j^b/\partial\alpha + \sum_{k=1,2}\partial D_j^b/\partial p_k \, dp_k/d\alpha + \sum_{k=1,2}\partial D_j^b/\partial W_k \, dW_k/d\alpha \qquad \text{(A.1)}$$

The change in black demand with respect to a change in lending standard is therefore the result of three effects:

- The partial equilibrium effect is $\partial D_j^b/\partial\alpha$ (the direct effect of lending standards);
- The general equilibrium effect going through prices is the second term $\sum_{k=1,2}\partial D_j^b/\partial p_k \, dp_k/d\alpha$;
- The general equilibrium effect going through neighborhoods' demographic composition

The response of prices $dp_k/d\alpha$ and Whites' demand to the lending constraint α remains to be determined.

At city equilibrium, demand equals supply in each neighborhood:

$$D_j(p_1,p_2,W_1,W_2;\alpha) = s_j(p_j), \qquad j=1,2 \qquad \text{(A.2)}$$

and the fraction of whites in neighborhood j equates the ratio of Whites' demand for housing to the overall supply of housing in neighborhood j:

$$W_j = D_j^w(p_1,p_2,W_1,W_2;\alpha)/s_j(p_j) \qquad \text{(A.3)}$$

The equilibrium is defined by equations (A.2) and (A.3). Ouazad and Rancière (2013) show that the equilibrium is unique when there are no preferences for same-race neighbors. With preferences for same-race neighbors, there are typically multiple equilibria, as in Brock and Durlauf (2002).

From equation A1, A2, A3, price changes implied by lending standard changes are implicitly determined by the following two equations

$$\partial D_j/\partial\alpha + \sum_{k=1,2}\partial D_j/\partial p_k \, dp_k/d\alpha + \sum_{j=1,2}\partial D_j/\partial W_k \, dW_k/d\alpha = ds_j/dp_j,\ j=1,2 \qquad \text{(A.4)}$$

which determines price changes $dp_1/d\alpha$ and $dp_2/d\alpha$ when lending constraints are relaxed. In the specific case of inelastic demand and no preference for white neighbors, we obtain simple comparative statics: a larger response of the demand for neighborhood 1 leads to a larger increase in the price of neighborhood 1.

Net Impact of Lending Standards on Black Households' Demand: Comparative Statics

Key structural parameters that pin down the impact of lending standards on segregation are the elasticity of black (resp., white) demand with respect to lending standards η_α^b (resp., η_α^w) at given prices, i.e. without accounting for the general equilibrium effect of lending standards on prices, the elasticity of prices with respect to lending standards $\eta_\alpha^{p_1}$ et $\eta_\alpha^{p_2}$, i.e. the general equilibrium effect of lending standards on prices, and the elasticity of black (resp., white) demand with respect to the price of housing in neighborhoods 1 and 2, $\eta_{p_1}^b$ and $\eta_{p_2}^b$, (resp. $\eta_{p_1}^w$ and $\eta_{p_2}^w$) i.e. the usual own-price elasticity of demand for housing.

A convenient assumption for the theoretical analysis is that black households do not have a particular taste or distate for white neighbors, formally $\partial D_j^w/\partial W_k=0$. None of the results we present rely on such assumption, but it leads to the following simple proposition: black demand for neighborhood 1 declines when credit supply increases whenever general equilibrium effects dominate partial equilibrium effects.

Proposition A1. *A relaxation of borrowing constraints leads to a net decline in black households' demand for the white-dominated neighborhood 1 whenever the elasticity η_α^b of black demand for neighborhood 1 with respect to lending standards α is lower than an increasing function of the price elasticity of black demand for neighborhood 1 (η_{p_1}), an increasing function of the elasticity of neighborhood 1's price with respect to lending standards α, and a decreasing function of the elasticity of neighborhood 2's price with respect to lending standards α:*

$$\eta_\alpha^b \leq \eta_{p_1}^b\left(\eta_\alpha^{p_1} - \frac{p_1}{p_2}\eta_\alpha^{p_2}\right) \qquad \textit{(A.5)}$$

The proof follows from writing that $dD_j^b/d\alpha < 0$ and replacing the expression for $dD_j^b/d\alpha$ coming from (A.1). The proposition formalizes the intuition that credit supply leads to a *decline* in black mobility towards the white neighborhood whenever price growth in neighborhood 1 offsets the beneficial effects of the relaxation of borrowing constraints. This is what inequality (A.5) states.

In particular, the following proposition shows that either a strongly inelastic housing supply in neighborhood 1, or a strong preference of white households for white neighbors, lead to a large values of the elasticity $\eta_{\alpha}^{p_1}$ of prices with respect to lending standards α.

Proposition A2. *Neighborhood 1's price elasticity* $\eta_{\alpha}^{p_1}$ *is (a) a decreasing function of neighborhood 1's housing supply elasticity, (b) an increasing function of neighborhood 2's housing supply elasticity and (c) an increasing function of white households' demand elasticity with respect to the fraction of Whites in neighborhood 1 – i.e., white households' preferences.*

Proof. The formula for the price change $dp_1/d\alpha$ in neighborhood 1 is obtained by solving the system of equations (A.4). Elementary calculations lead to the following equality relating the price elasticity $\eta_{\alpha}^{p_1}$ to the other elasticities of the model:

$$\eta_{\alpha}^{p_1} = \frac{-\left(\eta_{p_2}^{2} - \varepsilon_{p_2}^{2}\right)\left(\eta_{\alpha} + \left(\eta_{W_1} - \frac{W_1}{W_2}\eta_{W_2}\right)\eta_{\alpha}^{W_1}\right) + \eta_{p_2}(\eta_{\alpha}^{2} + \left(\eta_{W_1}^{2} - \frac{W_1}{W_2}\eta_{W_2}^{2}\right)\eta_{\alpha}^{W_1})}{(\eta_{p_1} - \varepsilon_1)(\eta_{p_2}^{2} - \varepsilon_2) - \eta_{p_1}^{2}\eta_{p_2}^{1}}$$

which implies the three points (a), (b), and (c) of the proposition: inelastic neighborhoods see larger price increases ((a) and (b)), and Whites' preferences for same-race neighbors leads to greater price increases (c). □