CREDIT STANDARDS AND SEGREGATION

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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 to 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 racially mixed neighborhoods. Without such a credit supply shock, black households would have had between 2.3 and 5.1 percentage points more white neighbors in 2010.

I. Introduction

THE availability and affordability of mortgage credit is a key determinant of housing choices. Large aggregate changes in mortgage lending standards 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, & Rueben, 2004; Cutler, Glaeser, & Vigdor, 1999), 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 & Yinger, 2002), and thus they 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 of relocating to more racially mixed neighborhoods but also neighborhoods with a comparable racial mix but 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 versus 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.

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. For each metropolitan area, we build mortgage loan approval rates and median loan-to-income (LTI) ratios for the credit boom of 2000 to 2006 and metrolevel measures of racial segregation and interracial exposure across census tracts in 2000 and 2010.

OLS estimation results show that in metropolitan areas that experienced larger increases in loan-to-income (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 more slowly 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 conditions—reflect both supply and demand factors.

This paper's identification strategy relies on instruments that identify the impact of the relaxation of lending standards, that is, the impact of the supply of credit separate from the impact of changes in credit demand. The instrumental variables are measures of banks' liquidity conditions at the metropolitan level in the 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

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[1] This paper's focus on metropolitan-level changes in segregation due to market forces is key: there is a literature on discrimination in mortgage lending and redlining (Ross & 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 & Vigdor, 2012).

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

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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 indeed suggest that metropolitan areas with a low level of bank liquidity over 1990 to 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. Moreover, cross-sectional regressions do reflect a catch-up effect: the cross-sectional relationship between lending standards and 1990–1994 liquidity turns from positive in 2000 to mostly flat in 2006. Bank liquidity in the early 1990s is thus a strong predictor of future changes in approval rates and in originations' loan-to-income ratio during the boom, yet bank liquidity in the early 1990s is not significantly correlated with observable factors affecting mortgage credit demand; most important, it is not significantly correlated with Hispanic inflows and income changes on average and by racial group. In addition, we find that our instruments do not display significant correlations 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 past decade (Glaeser & Vigdor, 2012), our results suggest that the increased supply of credit slowed the racial integration of cities.

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 fostering significant white mobility out of both racially mixed and mostly black Census tracts—tracts with between 10% and 60% black population and tracts with at least 60% black population—and into minority black Census tracts (tracts with less than 10% of black population). This mobility pattern with outflows starting at about 10% to 15% of black population is consistent with empirical estimates of a tipping point model (Card, Mas, & Rothstein, 2008). In addition, we find that lending standards relaxation led to significant black mobility into racially mixed tracts, which were experiencing white 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 toward 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.

Such microlevel 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 (2010), 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 II discusses related literature. Section III 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 IV presents the instrumental variables strategy and the paper's main results and discusses its robustness. Section V identifies economic mechanisms that are consistent with a simple model of neighborhood choice with credit constraints. Section VI concludes.

II. Related Literature

In 2000, a majority of urban blacks lived in highly segregated neighborhoods (Massey, 2004), and evidence suggests that racial segregation has negative impacts on black welfare: on education (Card & Rothstein, 2007), crime (Weiner, Lutz, & Ludwig, 2009), and black well-being (Massey, Condran, & Denton 1987). Analysis of racial segregation in the first half of the twentieth century stressed the importance of white households' collective action, including land use regulations and racial covenants, in shaping racial segregation (Cutler & Glaeser, 1997). Literature on racial segregation in the second half of the twentieth century models segregation as a market equilibrium, where price differences across white and minority neighborhoods reflect local amenities (Epple & Sieg, 1999), differences in households' income and education (Benabou, 1996), and preferences for same-race neighbors (Krysan & Farley, 2002). The white flight away from black neighborhoods

(Schelling, 1971) was facilitated by declining transportation costs (Baum-Snow, 2007) and inequalities in the quality of public services and education (Boustan, 2010). This paper's main goal is to contribute to this literature by understanding whether large changes in mortgage credit markets hinder or facilitate 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 creditworthiness covariates (Munnell et al., 1996), geographic controls (Ross & Tootell, 2004), and randomized experiments (Ross & Yinger, 2002). This paper stresses the possibility of large impacts of mortgage lending standards on racial segregation even absent significant racial discrimination in mortgage lending.

Evidence suggests that the increase in housing prices during the credit boom was accompanied by a large increase in the dispersion of prices across neighborhoods and substantial population flows across neighborhoods (Guerrieri, Hartley, & Hurst 2012). 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. Appendix section 4 formalizes this point in a simple model of location choice with borrowing constraints.

Identifying the impact of changes in lending standards 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 metrolevel racial segregation. Recent literature has shown that credit supply is responsible for a large share of the rise in leverage and approval rates during the credit boom. This literature includes Mian and Sufi (2009), Favara and Imbs (2015), Keys et al. (2010), Purnanandam (2011), and Adelino, Schoar, and Severino (2012).

III. Data Set and Descriptive Evidence

A. Data Sources

We use mortgage data for the years 1995 to 2007 compiled in accordance with the Home Mortgage Disclosure Act (HMDA), which mandates reporting by most depository and nondepository lending institutions. HMDA disclosure requirements thus apply to more than 90% of all mortgage applications and originations, and, for each mortgage, lender report of the loan amount, the applicant's income, the applicant's race and gender, 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 Massey, White, and Phua (1996) and Cutler et al. (1999). We equate "metropolitan areas" with the widely used Core Based Statistical Areas (CBSAs) in 2003 borders.[4] CBSAs encompass both metropolitan statistical areas (MSAs) and "micropolitan" statistical areas (μSAs).[5]

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 IV, 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.

B. Racial Segregation from 2000 to 2010

From the many available segregation measures (Massey & Denton, 1988) we choose the isolation and exposure indices, which have been extensively used in the literature (Cutler et al., 1999).[6] Here *isolation* is defined as the average fraction of neighbors of the same race in the average census tract of whites, blacks, or Hispanics;[7] thus, the isolation of whites is the average fraction of white neighbors for white households.

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

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

where $blacks_{k,j}$ is total black population in Census tract j of metropolitan area k. In the case of two racial groups, one

[4] We build 2000 tract to 2003 CBSA crosswalks to keep consistent metropolitan area borders across time.

[5] 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 (more than 10,000 inhabitants). Our metropolitan areas include both MSAs and μSAs.

[6] The isolation index is a particularly relevant measure as it is a measure of the contextual effect of neighbors on outcomes in linear peer effects specifications (Manski, 1993).

[7] Results for Asians are available on request.

group's isolation increases as exposure to the 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 & 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% 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.

C. *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). The median LTI ratio captures the extent to which a typical borrower can leverage her income.

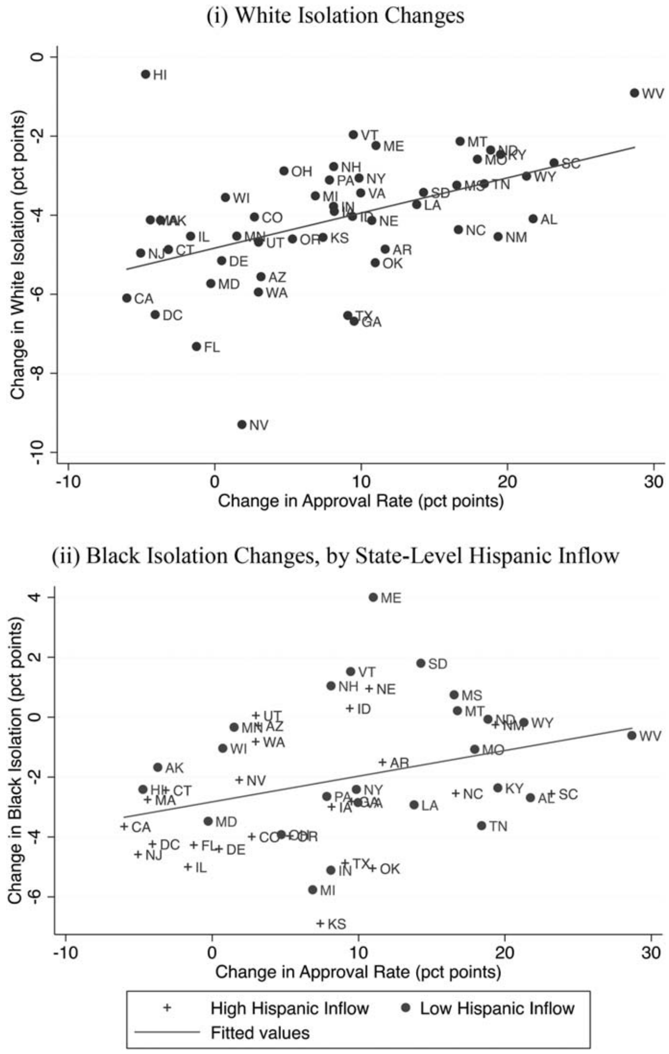
The median LTI for the entire population of mortgage originations increased from 1.89 to 2.3 during this period, with similar upward trends for the three major racial groups: an increase of 0.40 for white borrowers, 0.42 for black borrowers, and 0.41 for Hispanic borrowers.[8] Whereas the loan-to-income ratio increased dramatically during the boom (through 2006), the average LTV ratio showed little movement until 2006 (Gelain, Lansing, & Mendicino, 2012). The LTV ratio, however, misses the rise in multiple mortgage originations for the same property, which the combined loan-to-value ratio (CLTV) measures (Keys et al., 2013). Both the LTI and the CLTV rose substantially between 2000 and 2006.[9] The CLTV is hard to obtain at the metropolitan area level in this paper's years of observation. The rise of the LTI, correlated with the combined LTV ratio at the national level, thus seems to be a good indicator of the decline in underwriting standards.

Approval rates for mortgages increased significantly during the boom: from 70.1% to 84.2%. 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. These changes may be indicative of looser lending standards or of changes in the demand for housing or for credit. Section IV uses an instrumental variable strategy to disentangle these determinants of the change in credit market equilibrium conditions.

[8] Descriptive statistics are presented in the appendix (table A2).

[9] For prime fixed-rate mortgages, the CLTV ratio went from 74.5% in 2000 to 83.5% in 2006, with larger increases for adjustable-rate mortgages (Belsky, Herbert, & Molinsky, 2014).

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



States with high Hispanic inflows: the fraction of Hispanics grew by more than the median increase across states (+2.66 percentage points).

D. *OLS Estimates*

At the state level, figure 1 (top) displays a positive correlation between white isolation across urban Census tracts and the growth of mortgage approval rates. Increases in black and Hispanic segregation at the metropolitan area level are also positively correlated with 2000–2006 increases in approval rates (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:

$$\Delta Segregation_k = \gamma_{Approval}\Delta Approval\ Rate_k + \alpha_{Approval}\Delta Demographics_k + X_k\beta_{Approval} + State_{s(k)} + \varepsilon_k, \quad (1)$$

TABLE 1.—OLS REGRESSION OF CHANGES IN SEGREGATION ON CHANGES IN APPROVAL RATES AND LTI

	(1) Δ Black Isolation	(2) Δ Black Isolation	(3) Δ Black Isolation	(4) Δ Black Exposure to Whites	(5) Δ Black Exposure 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^2	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^2	0.619	0.657	0.721	0.673	0.892
	Δ Hispanic Isolation	Δ Hispanic Isolation	Δ Hispanic Isolation	Δ Hispanic Exposure to Whites	Δ Hispanic Exposure 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^2	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^2	0.744	0.782	0.845	0.790	0.740
	Δ White Isolation	Δ White Isolation	Δ White Isolation	Δ White Exposure to Hispanics	Δ White Exposure 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^2	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^2	0.926	0.929	0.957	0.948	0.773
Observations	939	939	939	939	939
Demographics	Yes	Yes	Yes	Yes	Yes
State effect	Yes	Yes	Yes	Yes	Yes
Additional variables	No	Yes	Yes	Yes	Yes
Income and price	No	No	Yes	No	No

This table presents regressions of changes in segregation on changes in credit conditions. Standard errors clustered at the state level. See data appendix section 1 for the list of additional control variables. Significant at $^{**}p < 0.01$, $^{*}p < 0.05$, $+\ p < 0.1$.

$$\Delta Segregation_k = \gamma_{LTI}\Delta Loan-to-Income_k + \alpha_{LTI}\Delta Demographics_k + X_k\beta_{LTI} + State_{s(k)} + \varepsilon_k, \quad (2)$$

where k indexes metropolitan areas (MSAs or μSAs) and $s(k)$ is the state of metropolitan area k. We use ΔSegregation$_k$ 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 IIIB for blacks, whites, and Hispanics. The term X_k is a set of observable controls, and $State_{s(k)}$ is a state fixed effect. The ε_k term is the residual clustered at the state level. Finally, $\Delta Demographics_k$ 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 (bottom), which shows states with a large increase in Hispanic population (above the median increase across states) as dots and states with a small increase in Hispanic population as crosses. The linear positive correlation between black isolation and changes in approval rates holds on both subsets.[10]

The OLS estimates γ_{LTI} and $\gamma_{Approval}$ are not causal; rather, they provide evidence of an economically and statistically significant correlation between changes in equilibrium credit conditions and changes in segregation. The possibility of a causal interpretation is presented in section D.

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 is 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

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

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 variables controlling for mortgage credit risk[11] 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 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 or changes in households' income, we incorporate in this column 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 panels of table 1 present 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 and OLS biases. OLS estimates of the impact of approval rate and LTI ratio changes on segregation simply correlate approval rate and LTI ratio changes with changes in segregation. Thus, OLS estimates (table 1 of the paper) are confounded by unobservable demand shifters and will be biased estimates of the impact of banks' lending policies on segregation. When borrowers' qualifications improve, approval rates typically increase; there is thus a positive correlation between borrowers' unobservable qualifications on the one hand and approval rate and LTI ratio changes on the other hand. The direction of the bias on OLS estimates crucially depends on the impact of borrowers' unobserved qualifications on racial segregation.

To determine the sign of the bias, let us consider the case in which an increase in black households' economic health translates into a higher loan-to-income ratio and simultaneously raises black demand for less segregated areas. The impact of borrowers' unobserved qualifications on racial segregation is negative. In that case, unobserved demand characteristics would bias downward the effect of a change in lending standards on black isolation. This downward bias, caused by black households' unobservables, could even lead to an estimated effect with an opposite sign and yield the prediction that banks' relaxation of lending standards reduces black isolation. However, this example is not sufficient to sign the overall direction of the OLS bias. An upward bias, caused by white households' unobservables, could also arise if improving white households' economic health translates into a higher LTI ratio, and simultaneously raise white demand for more racially homogeneous neighborhoods. The relative magnitude of these two unobservable demand shifters pins down the sign of the bias on the OLS estimate.

Borrowers' qualifications depend on a range of individual characteristics that are typically not observed for each race and for each of our 939 metropolitan areas, including households' assets and liabilities, income, and default risk. Nevertheless, the Current Population Survey includes wage income changes by race, from 2000 to 2006, for a subset of 230 metropolitan statistical areas (MSAS). Table A3 displays a positive significant correlation between black income growth and LTI ratio growth (+0.369), while the correlation between white income growth and LTI ratio growth is about three times smaller (+0.109). This suggests that the downward bias caused by black demand shifters may more than offset the upward bias caused by white demand shifters. For the LTI ratio changes, therefore, the OLS estimate will likely be a downward-biased estimate of the impact of banks' lending standards on black isolation.

Altogether the 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.

IV. Identification Strategy and Results

A. *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—before the mortgage credit boom of 2000 to 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 LTI 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.

The first such transformation favoring the rise in securitization of the industry was the Federal Housing Enterprises

[11] See appendix section 1 for the definition of the variables.

Financial Safety and Soundness Act of 1992.[12] This legislation established housing goals for the government-sponsored enterprises (GSEs, which include 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 to both the GSEs and banks holding the mortgage-backed securities (MBS) issued by GSEs.[13] This reduced "capital charges" treatment for the GSEs overturned previous recommendations from the U.S. Treasury that GSEs should actually increase their capital in order to comply with the Basel Committee's risk-based capital rules (U.S. Treasury, 1990). 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 provided loan "warehousing" and ultimately securitization via private entities (Levitin & Wachter, 2012). The development of ABCP conduits was boosted by new accounting rules (Basel agreements, 1992) 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 (Acharya et al., 2010). The ABCP market grew from about \$600 billion in 2000 to \$1.2 trillion in 2006, by which time it had become the largest U.S. money market instrument. Moreover, the share of mortgage loans in new ABCP issuances increased from 36% in 2000 to 70% in 2006 (Adrian & Shin, 2010). These two structural transformations were the main contributors to the mortgage industry's change from a predominantly "originate and hold" model to a 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% (U.S. Flow of Funds accounts, table L.218).

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 LTI 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

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

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.[14] 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 U.S. dollars) in the area,

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

and similarly for securitizability. Here 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.

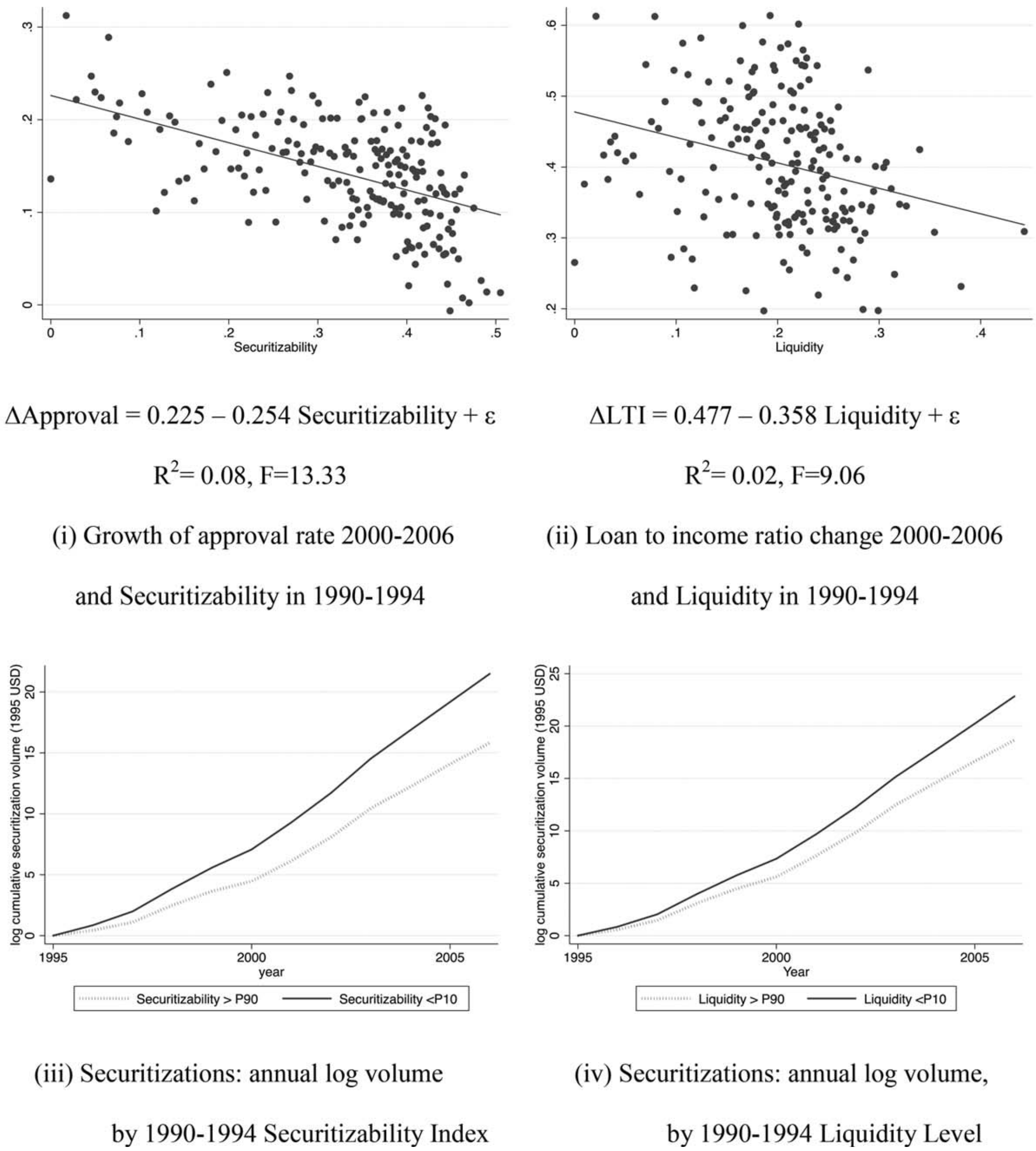
Finally, we average the metropolitan area liquid asset ratios (resp., securitizability) over the 1990–1994 period and use them as instruments for the 2000–2006 growth in the median LTI ratios (resp., the approval rate). There is significant independent variation of our liquidity and securi-

[12] The FHEFSA Act was not implemented immediately. The act set 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 to 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 to 1994 to construct our instrument.

[13] 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, & Van Nieuwerburgh, 2011).

[14] Loan portfolios are broken down into six types of loans (Loutskina, 2011): (a) home mortgages, (b) multifamily residential mortgages, (c) commercial mortgages, (d) consumer credit, (e) business loans not secured by real estate (commercial and industrial loans), and (f) farm mortgages. Securitizability measures are based on the U.S. Flow of Funds accounts, and individual bank-level loan data are from each bank's Report of Income and Condition.

FIGURE 2.—LENDING STANDARDS, LIQUIDITY MEASURES, AND SECURITIZATIONS



This graph presents the log increase in securitization for metropolitan areas with securitizability (left) and liquidity (right) below the 10th percentile of the metropolitan area distribution of the liquidity measure, above the 90th percentile.

tizability 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. 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. Figure 2 (top) shows a negative correlation between initial liquidity measures and 2000–2006 changes in lending standards. Figure 2i shows that low-securitizability metropolitan areas experienced a greater increase in approval rates during the boom. Figure 2ii shows similar (albeit less strong) results regarding the link between bank liquidity and LTI ratios.

Furthermore, the negative relationship between lending standard changes and initial liquidity likely corresponds to a catch-up effect. Figure A1 plots the cross-sectional relationship between the approval rate and the 1990–1994 mea-

TABLE 2.—FIRST STAGE: GROWTH OF APPROVAL RATES AND LOAN PORTFOLIO SECURITIZABILITY

	(1) ΔApproval	(2) ΔApproval	(3) ΔApproval	(4) ΔApproval	(5) Δlog Securitization
Securitizability	−0.225**	−0.155**			−0.287**
	(0.062)	(0.028)			(0.091)
Securitizability			−0.169**		
(limited market share)			(0.027)		
Securitizability				−0.352**	
(computed without mortgages)				(0.067)	
Observations	939	939	939	939	939
R^2	0.175	0.625	0.629	0.627	0.483
Demographics	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-statistic for H_0: Securitizability Coefficient $= 0$	13.07	29.81	38.31	28.08	10.03

Limited market share: banks with less than 25% of market share.

TABLE 3.—FIRST STAGE: GROWTH IN LOAN-TO-INCOME RATIO AND BANK LIQUIDITY

Variables	(1) ΔLTI	(2) ΔLTI	(3) ΔLTI	(4) ΔLTI	(5) Δlog Securitization
Liquidity	−0.329**	−0.196**			−0.509**
	(0.113)	(0.069)			(0.174)
Liquidity			–0.188*		
(limited market share)			(0.079)		
Liquidity				−0.118*	
(computed without mortgages)				(0.056)	
Observations	939	939	939	939	939
R^2	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-statistic for H_0: Liquidity Coefficient: $= 0$	8.50	8.07	5.63	4.53	8.57

Limited market share: banks with less than 25% of market share.

sure of securitizability for 2000 (white dots) and 2006 (black dots). There is a strongly positive cross-sectional relationship between 1990–1994 securitizability and 2000 mortgage approval rates: banks with the lowest (resp., highest) level of securitizability approved 40% of their applications (resp., 57%). In 2006, the relationship becomes almost flat as the slope of the approval rate-securitizability relationship declines substantially, from 0.35 to 0.06.[15] Similarly, figure A2 plots the cross-sectional relationship between LTI ratio and the 1990–1994 measure of liquidity in 2000 and 2006. The figure and the associated linear regression results show a positive and significant relationship between bank liquidity and the LTI ratio in 2000 (+0.417), which becomes nonsignificant in 2006 (+0.198).

LTI growth and approval rate increases are significantly and positively correlated with the growth of securitization volumes. Figure 2 (bottom) 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 to 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 13.07 (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. The corresponding Cragg–Donald statistic (described later) provides an estimate of finite sample IV bias. 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 performs within-state estimation using state effects. Results are robust to the inclusion of state effects as coefficients remain significant at 1% and not statistically different from column 1. The last column confirms that metropolitan areas with a low level of liquidity or

[15] Table A4 in the appendix shows the year-by-year estimates of the linear relationships between approval rate and securitizability, and between LTI and liquidity, between 2000 and 2006. The cross-sectional slope coefficient declines from 2000 to 2006.

securitizability experienced a higher growth of log securitization volume.

Three potential concerns affect the validity of our instrument. The first concern is that banks' liquidity and loan portfolio securitizability from 1990 to 1994 may reflect fundamental demand characteristics rather than being determinants of future credit supply shifts. In particular, loan 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 using two alternative measures of liquidity and securitizability. First, we focus on the liquidity of banks for which the metropolitan area accounts for less than 25% of the bank's originations. For these banks, metropolitan area demand is unlikely to have much effect on liquidity and loan portfolio composition. Coefficients estimated using these measures (column 3 of tables 2 and 3) are similar and not statistically different from those in column 1 of these tables. Second, liquidity and securitizability excluding residential mortgages are less likely to reflect households' preferences for homeownership or credit.[16] Results using this alternative measure of securitizability are similar.

The second concern is that our instrument may be 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 A5.1, 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 noninstrumented changes in the LTI ratio and approval rate are correlated with several of these observable factors.

The third concern is that instruments might be correlated with other structural determinants of racial segregation that could vary over time or whose impact on segregation varies over time. Table A5.2, column 3, indeed shows that the change in income inequality (Gini 2010 minus Gini 2000), the burglary rate, the homicide rate,[17] 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. The lack of significant correlation of our instruments with measures of overall income or racial inequality (the black-white income gap) is important. It implies, in particular, that the finding that reduced racial inequality may lead to increased racial segregation (Bayer, Fang, McMillan, 2005) is unlikely to drive our IV estimation results. Column 1 (resp., column 2) shows also that coefficients of the regressions of the same potential confounders on our liquidity instrument (resp., our securitizability instrument) are not significant.

We perform a further test to compare trends in three observable characteristics of metropolitan areas with more or less liquid banks (resp., securitizable banks): crime rate, wages (from income tax returns), and the fraction of employment in manufacturing (a proxy for the fraction of employment in the tradable sector). We compare trends in each of these three variables between high liquidity and low liquidity (resp., securitizability) metropolitan areas in figures A3 and A4 in the following way. First the sample is split into two samples of equal sizes. High-liquidity (resp., high-securitizability) metropolitan areas have a liquidity ratio in 1990 higher than the median liquidity ratio. Each of the three dependent variables (manufacturing share, income, crime rate) is regressed on a set of year dummies, and a set of year dummies interacted with a indicator variable for whether the metropolitan area had high liquidity (resp., securitizability) in 1990. Any significant coefficient for any of these year dummies interacted with the high-liquidity indicator would be a sign that the trends in manufacturing share, income, and crime rate differ across the two groups. Results of the six regressions are presented in figures A3 and A4, panels a to c, with corresponding 95% confidence bands for each point estimate. Results suggest that trends of crime, wage, and share of employment in manufacturing are not significantly different across high and low securitizability and high- and low-liquidity areas.

B. *Instrumental Variable Estimates*

Our IV estimates are presented in table 4. The change in LTI 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% to 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

[16] See the appendix data section 1 for the definition of the alternative measures.

[17] The literature suggests that racial segregation and crime are strongly related (Krivo & Peterson, 1996).

TABLE 4.—IV ESTIMATES OF THE IMPACT OF LENDING STANDARDS ON SEGREGATION

	(1)	(2) Instrumental Variable Regression	(3)		(4) Reduced Form
Dependent Variable:	ΔBlack Isolation	ΔBlack Exposure to Whites	ΔBlack Exposure 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^2	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^2	0.15
First-stage F-statistic	8.501	8.501	8.501		
Dependent Variable:	Δ Hispanic Isolation	Δ Hispanic Exposure to Whites	Δ Hispanic Exposure 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^2	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^2	0.65
First-stage F-statistic	8.501	8.501	8.501		
Dependent Variable:	ΔWhite Isolation	ΔWhite Exposure to Hispanics	ΔWhite Exposure 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^2	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^2	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 parentheses. Significance levels: $**p < 0.01$, $*p < 0.05$, $+\ p < 0.1$.

do not present significant evidence that instrument weakness would substantially alter our main conclusions.

In table 4, the impact of the change in approval rate on the change in black isolation is significant, strong, and positive. The IV coefficient of the effect of a change in approval rate on the change in black isolation (18.1**) is the mirror image of the coefficient of the change in black exposure to whites (−17.76**), while the effect on black exposure to Hispanics is marginal (−0.821) and insignificant. This implies that the increase in black isolation is entirely due to the decrease in the exposure of blacks to whites. A 1-standard deviation increase in the approval rate, which is about a 10 percentage point increase, is associated with a 1.8 percentage point increase in black isolation due to a decline (also of 1.8 percentage points) in black exposure to whites; in this case, there is also a (nonsignificant) decline of 0.1 percentage point 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 over 1990 to 1994 witnessed a significantly greater 2000–2010 decline in black isolation.

On average, black exposure to white neighbors declined from 33.2% in 2000 to 32.7% 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 discussion in section IIID suggests that OLS estimates are likely downward-biased estimates of the impact of banks' lending standards on black isolation. Indeed, we find that the difference between the IV and the OLS estimates of the impact of LTI on black isolation is positive, large, and significant (20.24 for the IV versus 3.29 for the OLS). In the case of approval rate, the difference is positive but smaller and not significant (18.1 for the IV versus 11.9 for the OLS).

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 the 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.

Finally, one may wonder whether the mechanism at play was already present before the credit boom. Table A6 of the appendix, which replicates the paper's main IV estimation for the period 1990 to 2000, suggests that the identification strategy specifically captures the effect of 2000 to 2006 changes in lending standards on racial segregation.

The role of housing supply elasticity. Table A8 in the appendix presents IV estimates in which changes in lending standards are interacted with metrolevel housing supply elasticity (Saiz, 2012). 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. For both approval rate (upper panel) and LTI (lower panel), 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 noninteracted coefficients for changes in lending standards keep the same sign and significance as in our baseline IV estimates. For the metropolitan area with average elasticity (2.49), a 10 percentage point increase in the approval rate leads to a 0.86 percentage point increase in black isolation (first line, upper panel), and a 0.4 increase in the median loan-to-income ratio leads to a $11.74 \times 0.4 = 4.7$ percentage point 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 percentage point increase in approval rate leads to a $0.86 - 0.10 \times 4.8 \times (1.65 - 2.49) = 1.26$ percentage point increase in black isolation, an effect that is 47% larger than for the average metropolitan area.

The role of the housing bust. An important issue deals with the timing of the measurement of changes in segregation. Tract-level U.S. Census data are 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 postboom events between 2007 and 2010 (foreclosures, labor market shocks, and other factors) may be correlated with credit conditions during the boom years. We tackle this issue in two ways: by controlling directly for foreclosures and then by using two additional data sets where the timing of demographic data more closely matches the timing of credit data: 2006–2010 American Community Survey Data and *annual* school-level demographics (U.S. Department of Education).

Results on the three data sets, presented in the appendix's tables A9 to A11, confirm our main findings: the effects of foreclosure appear largely orthogonal to the segregative effect of lending standards' relaxation; estimates using ACS data for 2006 to 2010 are similar to estimates using 2010 census; and estimates using 2000 to 2005 school segregation are similar to estimates using 2000 to 2010 urban segregation.

Finally, two additional other robustness checks, dealing with multiple mortgage applications and the endogeneity of metrolevel migration flows, are presented in the appendix.

V. Inspecting the Mechanism: Microevidence of Population Flows

A. *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, and, in fact, almost entirely because of such effect, when changes in lending standards are measured using approval rates. In this section, we explore how Census tract data on the mobility of blacks and whites can shed light on this result. In the section 3.a in the online appendix, we expand the Census tract analysis to include the role of Hispanic mobility.

Lower black exposure to whites could be due to a combination of white households' mobility out of racially mixed neighborhoods and into predominantly white neighborhoods or black households' mobility out of racially mixed neighborhoods and into predominantly black neighborhoods, or both.[18]

We tell these two different stories apart by building a neighborhood data set: a matched longitudinal census tract level data set based on the 2000 and 2010 U.S. Censuses. 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. 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 et al. (2008), as the change in white population as a fraction of the tract's total population in 2000. Census tracts are grouped into $K = 20$ bins (indexed by k), ranked according to their fraction of black residents in the 2000 Census: the 1st and twentieth bins 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 Census tract demographic changes in the following way. First, we sort metropolitan areas in two groups: 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—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 twenty 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

$$\begin{aligned}\Delta White\,Pop_c = \sum_{k=1}^{K} \big(&\delta_{k,high} Bin_k \times High\Delta \widehat{Approval}_{metro(c)} \\ &+ \delta_{k,low} Bin_k \times Low\,\Delta \widehat{Approval}_{metro(c)}\big) \\ &+ \alpha \Delta Demographics_{metro(c)} \\ &+ State_{s(c)} + \varepsilon_c, \qquad (3)\end{aligned}$$

includes state fixed effects and metropolitan-level measures of demographic changes by race. Standard errors are clustered at the state level. State fixed effects have a 0 average and metro-area demographic changes are cross-sectionally demeaned; hence, 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 et al. (2008), conditioning on metropolitan area demographic changes by race in the specification is key. These demographic variables not only capture population flows, by race, into and out of the metro area. They also control for internal population dynamics due to birth rate/death rate differentials. Providing that these differentials vary between races and between metropolitan areas, but not by race across tracts within metropolitan areas, the associated demographic changes at the tract level are indeed captured by the metrolevel 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. 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 affected by changes in approval rates.

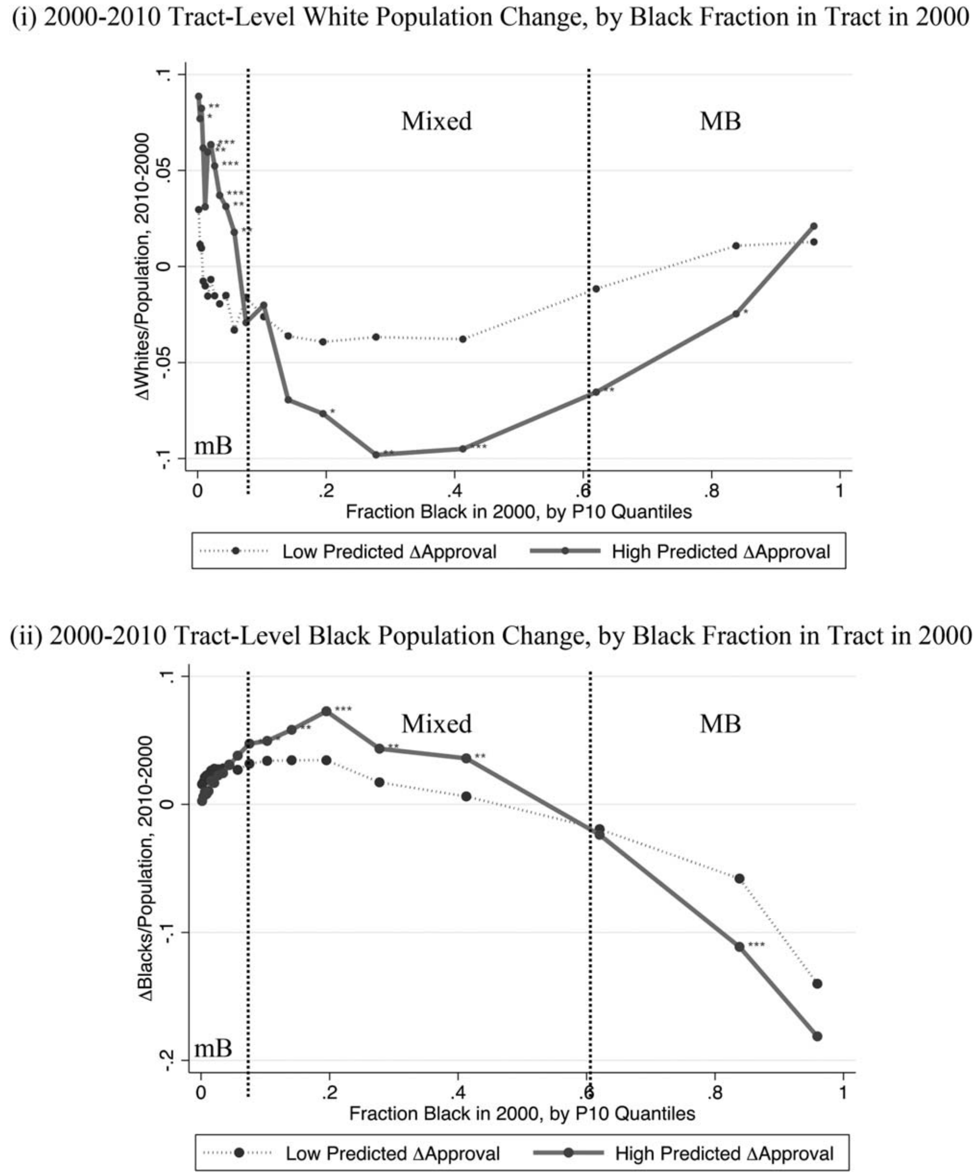
Figure 3i shows that white households moved into Census tracts with less than 7% to 8% blacks in 2000, but only in Census tracts with a high predicted increase in approval rates. In tracts with high-predicted approval rate increase, white inflows range from about 9% (in almost completely white tracts) to 2% (in tracts with about 6% black population). In sharp contrast, in metropolitan areas with low predicted approval rate increase, white households did not move into tracts with a similarly low fraction of black population. Racially mixed tracts—with an initial share of black population ranging from 10% to 60%—have all experienced white outflows. These outflows are about two to three times larger in tracts with a high predicted increase in approval rates (ranging from 7% to 10%). Similar white outflows occurred in predominantly black tracts (tracts with 60% of more blacks) at the exception of the very last bin of tracts, where there were almost no white residents in 2000.[19]

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 minority black neighborhoods (less than 10% black). These figures are consistent with the estimates of a tipping point model (Card et al., 2008), which predicts significant white outflows from tracts that have between 5% and 20% racial minorities (i.e., our estimate of around 10% is within these bounds). The novelty of this paper's results lies in the fact that these dynamics of segre-

[18] These neighborhoods include desirable neighborhoods with a significant fraction of college-educated black neighborhoods.

[19] In that bin, tracts are about 95% black.

FIGURE 3.—INFLOWS INTO CENSUS TRACTS, BY PERCENTAGE BLACK IN TRACT



Significant of ***1%, **5%, *10% of the *F*-test that the coefficient for high-liquidity metropolitan areas is equal to the coefficient for low-liquidity metropolitan areas. mB: fraction black below 10%. Mixed: between 10% and 60%. MB: above 60%.

gation have been considerably more pronounced in metropolitan areas that experienced a larger relaxation of lending standards during the credit boom.

Figure 3ii describes black inflows for the same Census tract bins as in panel i. We observe sizable 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%) in tracts with 20% of black population and high (low) predicted increase in approval rates. In tracts with an 80% initial share of black population, outflows are about twice larger (11.1 versus 5.8 percentage points) in Census tracts located in metro areas with a high-predicted approval rate increase.

Racially mixed neighborhoods (those with 10% to 60% black population) experienced white outflows and black inflows over the same time period. Furthermore, these black inflows and white outflows are of very similar magnitude, especially for neighborhoods with initially 15% to 20% of black population. This suggests that outgoing whites have been replaced by incoming blacks in those racially mixed neighborhoods; such a pattern is 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 authors, these population declines have been facilitated by declines in lending standards: in tracts with 80% (95%) blacks, population has shrunk by about 10% (15%) percent in areas with high predicted approval rate increases, but only by about 1% (10%) percent in areas with low predicted approval rate increases.

As a robustness test, we replicate the results discussed here by using propensity score matching estimates instead of OLS estimates. The methodology and the results are presented in the appendix. Matching on additional tract-level covariates yields similar estimates.

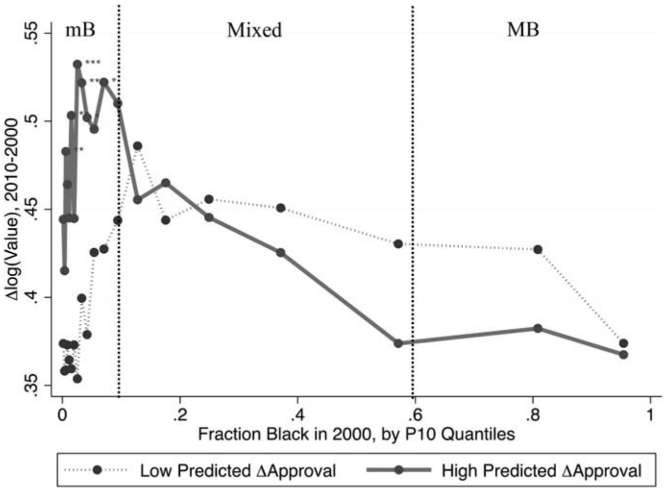
To sum up, the findings just described allow an interpretation of our metro-level findings (see section IV) as the result of two mobility trends, both significantly more pronounced in metropolitan areas with a larger predicted relaxation of lending standards. The first trend is that of black mobility 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 minority black neighborhoods. This second trend has contributed to the reduction of the exposure of blacks to whites. This microevidence 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.

B. *Explaining White and Black Mobility: Price Changes and General Equilibrium Effects*

Our tract-level results show that credit standards allowed significant black population mobility 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—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 toward mostly nonblack neighborhoods (fraction black lower than 10%), and that lending standards *did* make it easier for whites to move to those mostly white neighborhoods.

One potential explanation for the the lack of black mobility toward mostly nonblack neighborhoods is that 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 black households. Evidence of the impact of credit supply on county-level house prices is presented in Imbs and Favara (2015).

Figure 4.—Price Appreciation 2000–2008 and Fraction Black in 2000



Stars: High predicted ΔApproval: metropolitan level approval rates for mortgage applications increased by more than 13 percentage points. Standard errors clustered by state. mB: minority black tract, fraction black below 10%. Mixed: Mixed tract, fraction black between 10 and 60%. MB: Mostly black tract, fraction black above 60%. Significant at ***1%, **5%, *10%.

We present evidence of the tract-level mechanism using Census tract median house value data. Price data at the tract level are typically difficult to obtain for the entire universe of metropolitan areas. Detailed transaction-level data are usually available for metro-level subsamples. At the national level, the 2000 long-form version of the Census and the 2010 tract-level American Community Survey provide self-declared estimates of house values that can allow us to build a 2000–2010 longitudinal sample of tracts.

Figure 4 plots tract-level log house value appreciation between 2000 and 2010 for the twenty bins of tracts sorted by the fraction of blacks in the neighborhood. The graph was constructed as in the previous section, 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. Including the 2007–2010 postal code foreclosure rates per housing unit from RealtyTrac does not significantly affect estimates. 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%).

Mostly nonblack tracts (tracts with fewer than 10% of black population) saw significantly higher price increases in metropolitan areas with a higher-than-median predicted approval rates (greater than 13 percentage points) than in metropolitan areas with lower-than-median approval rates (lower than 13 percentage points) (figure 4). Price increase differences are both large—ranging between 6% and 18%—and almost all significant. A comparison of figure 3 (upper panel) and figure 4 shows that large price increases closely coincide with large white inflows in the same mostly nonblack tracts with high predicted increases in approval rate. In the rest of tracts, there 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 only 38% in tracts with 95.4% of blacks, a 15 percentage point difference.

In mostly nonblack tracts, figure A6 in the appendix suggests 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 a 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 nonblack areas is sufficiently low to generate the price effects illustrated in figure 4 and thus the segregative effects of credit supply.

The appendix's section 4 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.

C. Exposure to Poverty, Crime, and Dropout Rates

Welfare implications also depend on whether household mobility affected exposure to poverty, crime, and low-achieving students.

Figure A7, panels a and c, shows, perhaps unsurprisingly, that in the 2000 cross-section, poverty rates and high school dropout rates[20] increase with the fraction of black population. Combining this information with the mobility pattern shown in figure 3 indicates that lending standards' relaxation enabled black households' mobility out of neighborhoods with a poverty rate around 25% (resp., with dropout rates above 27%), and into neighborhoods with a poverty rate in the 10% to 15% range (resp., with a 15% to 20% dropout rate). Mortgage credit also allowed the mobility of white households from neighborhoods with a 14% to 23% poverty rate, and into neighborhoods with a poverty rate below 10%. The red line of the same figure A7(a) suggests that a neighborhood's ranking by its 2010 poverty level does not differ substantially from its ranking by 2000 poverty level.

Figure A7b focuses on the exposure to crime using the 2000 National Neighborhood Crime Study (NNCS). The slope of the fraction black-crime relationship is substantially larger for neighborhoods with less than 10% black population. Hence, white households moving from racially mixed neighborhoods toward mostly nonblack neighborhoods experience a larger reduction in their crime exposure than black households moving in to racially mixed neighborhoods.

[20] High school dropout definitions change in the 2000–2010 decade. Although 2010 high school dropout data are available, continuous school district-level time series data from 2000 to 2010 are not available.

VI. 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 on urban segregation. Although the downward trend in racial segregation continued during the first decade of the twenty first century, our instrumental variables estimates reveal that this decline has been noticeably slower in metropolitan areas where lending standards were most relaxed.

Census-tract-level results provide an understanding of the demographic and prices changes that underlie metropolitan level results. In areas where credit standards were most relaxed, white population declined in racially mixed neighborhoods and increased in mostly white neighborhoods as prices increased significantly more in such mostly nonblack neighborhoods. Credit standards did allow blacks to move from mostly black to racially mixed neighborhoods, but black exposure to whites did not increase substantially: white mobility toward mostly white neighborhoods dampened any such effect.

Results show that relaxing access to mortgage markets does not reduce such racial segregation; the deregulation of mortgage credit markets likely amplifies the effect of decentralized sorting on segregation. These results are important because of the negative effect of segregation on education (Card & Rothstein 2007; Hanushek, Kain, & Rivkin 2009), crime (Weiner et al., 2009), and black well-being (Massey et al., 1987).

The results of this paper also 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. Measures designed to address urban segregation will likely need to do more than simply expand access to mortgage credit or, more generally, opening up real estate or mortgage markets.

Our paper uses the more recent expansion of credit over 2000 to 2006 as a large-scale experiment to measure the impact of lending standards relaxation on segregation. Yet other significant episodes of mortgage credit expansion in the 1920s and in the postwar period occurred during a period of increasing racial segregation. Boustan (2010) provides evidence that postwar white flight from central cities contributed to the significant increase in racial segregation. While beyond the scope of this paper, establishing the role mortgage credit played in postwar white flight is an interesting avenue for future research.

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