The Voters' Veto: Local Racial Demographic Change and Exclusionary Behavior

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Abstract

As Western democracies diversify, racial threat risks activating exclusionary attitudes among members of the majority. While research has focused on support for immigration restrictions, local housing policy is the primary means of maintaining racial segregation in the US. Using referendum vote share from over 3,700 precincts in Los Angeles County, I show that an increase in the local non-white population leads to greater behavioral support for direct democratic control over affordable housing. I argue that this "voters' veto" is valued for excluding low-income, predominantly non-white residents from one's community. This effect exists in both majority white and non-white precincts, demonstrating that exclusionary reactions are not limited to racial threat among the majority. Instead, a shrinking white population may signal broadly shared concerns about neighborhood trajectory. Given local conditions can shape political beliefs, these findings provide a mechanism for the shift of white and non-white voters towards exclusionary political movements.

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The United States and other Western democracies are rapidly racially diversifying, largely due to immigration (Frey 2018). According to theories of racial threat, this increase in the multiracial minority population may present an economic, cultural, and political threat to the majority demographic group (Blalock 1967; Blumer 1958). Empirically, a growing minority population at the local and national level has been found to elicit exclusionary attitudes and behaviors among members of the majority racial group (see Craig, Rucker and Richeson (2018) and Kaufmann and Goodwin (2018) for review). However, three overlapping limitations have hampered the advancement of theory in this space.

First, researchers have largely focused on the exclusionary responses of the current majority group in these democracies — non-Hispanic whites (e.g., Abrajano and Hajnal 2015; Campbell, Wong and Citrin 2006; Fouka and Tabellini 2022; Outten et al. 2012; Reny and Newman 2018; Sahn Forthcoming). This emphasis on white voters risks oversimplifying emerging political cleavages into a majority-minority dichotomy and overlooking the motivations and behaviors of Black, Latino, Asian, and other non-white voters.[1] While white voters may be driven by racial threat, minority residents may also respond to the attendant changes that come with local demographic change, such as concerns over the social and economic stability of their community. Still, research which has examined the attitudes of minority residents has largely relied on static distributions of group size to understand intergroup conflict rather than the dynamic process of demographic change (e.g., Bobo and Hutchings 1996; Gay 2006; Oliver and Wong 2003).

Second, research on demographic change and racial threat has focused heavily on either exclusionary attitudes or on behaviors that only indirectly shape policy outcomes. Examples of these indirect behaviors include increased voter turnout and support for conservative candidates (e.g., Enos 2017; Newman, Shah and Collingwood 2018; Hill, Hopkins and Huber 2019). But neither of these dependent variables directly captures voters' behavioral support for exclusionary policy. This limitation likely stems from the fact that citizens rarely get to

[1] I use "Latino" to reference those identifying as of Hispanic, Latino, or Spanish origin via the Census.

vote directly on national exclusionary policy, such as immigration quotas (cf. Hainmueller and Hangartner 2013).

Third, studies measuring voters' policy preferences as the outcome tend to emphasize national policy (e.g., Enos 2014; Hopkins 2010). While immigration policy influences national demographics, a voter's local exposure to demographic change in the United States is heavily shaped by their municipality's zoning and land use policy. Voters and elected officials often use zoning practices to maintain economic segregation and therefore racial segregation (Trounstine 2018). From the enshrinement of single-family zoning to the obstruction of affordable housing, the practice of "exclusionary zoning" helps establish a local price floor that prevents the immigration of lower-income residents into the municipality (Einstein, Glick and Palmer 2020; Rothwell and Massey 2009; Trounstine 2020). Taken to an extreme, residents can stymie integration via direct democracy, where housing proposals and land use policies are placed on the local ballot (e.g., Hankinson 2018). In short, an array of local regulatory tools give voters substantial veto power over the construction of new housing, entrenching existing patterns of segregation.

Given the importance of measuring social context using the scale at which conflict occurs (Oliver and Mendelberg 2000; Oliver and Wong 2003), the effect of local demographic change on local behavior has been under-examined. Returning to the first limitation, studies which have looked at the relationship between local change and local exclusionary policy focus on institutional (i.e., citywide rather than individual) outcomes attributable to the preferences of white homeowners. For example, as southern Blacks moved into northern cities during the Great Migration, whites not only left central cities for the surrounding suburbs, but also used zoning to limit the ability of Black migrants to follow them (Sahn Forthcoming). Although increasing diversity is expected to have the largest effect in white communities with few minorities to begin with (Newman 2013), far less attention has been paid to the racially diverse communities experiencing more recent changes in their local demographics.

In this study, I address these three limitations by measuring the effect of local racial demographic change on local exclusionary behaviors across an array of racially diverse communities. As an exclusionary outcome, I use a California state law requiring that any proposal for new "low-rent" housing must win the support of a majority of voters via the municipal-level ballot.[2] Known as "Article 34," this policy transferred final decision making over affordable housing from the city council to the local electorate. Recognizing that placing this veto power in the hands of voters made it more difficult to build affordable housing, the California legislature has tried repeatedly to repeal Article 34 via statewide ballot referenda, with failed attempts in 1974, 1980, and 1993.[3] These repeal efforts provide an opportunity to observe voters' behavioral support for an exclusionary tool: direct veto power over new affordable housing.

Using electoral returns from 3,718 precincts across Los Angeles County in the 1993 repeal effort, I find that an increase in the local non-white population from 1980 to 1993 leads to higher precinct-level support for this "voters' veto" over affordable housing. I use two analytical strategies to assess this relationship. First, I show that a standard deviation increase in a precinct's non-white population increases support for the voters' veto by 4.5 percentage points ($d = .32$); similar effects are found when defining change as a proportion. Additionally, I show that this effect is not explained by alternative explanations, such as local changes in population density, residential churn, median household income, and percent foreign born.

Second, I find that this exclusionary response emerges in not only majority white, but also more racially diverse precincts, including majority Black and majority Latino precincts. This effect persists among white and non-white voters even when using Ecological Inference methods to estimate behavior by racial groups (King 1997). As individual-level evidence, I also use geocoded survey data to show that non-white respondents are more likely to support exclusionary housing policy when living in a ZIP code which has experienced a recent increase

[2] Though initially targeted towards public housing, "low-rent housing" has been interpreted to include any housing in which 50% or more of the units are government-subsidized.

[3] See Section A for the full text of Article 34 and the 1993 repeal ballot proposition.

in its non-white population.

Third, to account for the possibility that composition changes (i.e., new voters in each precinct) are driving this change, I also measure the effect of a precinct's spatial proximity to neighborhoods in which the non-white population increased most rapidly. Using only precincts which had experienced both minimal residential churn and changes in their own racial demographics over this period of time, a standard deviation increase in proximity to the nearest diversifying area leads to higher support for the voters' veto (mean +3.5 percentage points, $d = .24$). Again, I find this exclusionary response to nearby demographic change in both majority white and more racially diverse precincts.

Together, these findings advance our understanding of the political effects of local racial demographic change on voters' behavioral support for exclusionary policy. Exclusionary responses to demographic change are not limited to the members of the majority group via the mechanism of racial threat. Instead, an increase in the local non-white population may come with political and socioeconomic concerns shared across a wide swath of current residents. For instance, the exclusionary response observed in this study aligns with other documented efforts by middle-class majority Black neighborhoods to separate themselves from poorer Black communities and "fortify their neighborhoods against this encroachment" (Pattillo-McCoy 1999, p. 6). As middle-class minorities increasingly move to suburban communities (Badger, Bui and Gebeloff 2019; Orfield and Luce 2013), these findings suggest that both white and non-white voters will protect existing exclusionary zoning.

Beyond local policy, there is growing evidence that sizable shares of Latino, Asian, and even Black voters may be moving away from the Democratic Party (Fraga, Velez and West Forthcoming; Geiger and Reny 2024; Kao 2023; Sommer and Franco 2023), specifically in areas with high levels of immigration (Cai and Fessenden 2020). Given the ability of local conditions to shape social identities and even partisanship (e.g., Cramer 2016; Larsen et al. 2019; Ternullo 2024), these findings may provide a local mechanism for understanding the national shift of many non-white voters away from the Democratic Party and towards more exclusionary political movements. Finally, as the United States struggles with a deepening affordability crisis, these findings underscore the tension between integration and local control over housing in a diversifying society.

Direct Democracy as an Exclusionary Institution

In the United States, direct democracy has a long history of enabling voters to bypass legislative processes and make policy via ballot measures (Gerber 2011). Within the state and local context, direct democracy frequently targets housing and has been used to perpetuate segregation. In a 1964 statewide election, California voters in predominantly white cities near diversifying areas were more likely to support a ballot measure allowing property sellers, landlords, and real estate agents to racially discriminate (Reny and Newman 2018).

Once such *de jure* segregation was prohibited, voters began constructing economic barriers to integration (Trounstine 2018). By preventing the construction of more affordable multifamily housing, single-family zoning helps set a minimum price of residency. In turn, municipalities near center cities which experienced greater levels of mid-century Black migration were more likely to adopt single-family zoning (Sahn Forthcoming). This use of policy levers to economically "defend" one's community is collectively known as "exclusionary zoning" (Einstein, Glick and Palmer 2020).

In this study, I focus on voter support for repealing one of these exclusionary tools: direct democratic control over affordable housing. While the institution appears race-neutral, the context of its implementation underscores the tension between racial demographic change and support for voters' veto power. Following World War II, Congress passed the Housing Act of 1949, promising "a decent home and suitable living environment for every American family" by building 810,000 units in six years. Simultaneously, the Second Great Migration was accelerating, with ultimately more than 5 million Blacks leaving the Southeast and moving to the Northeast, Midwest, and West Coast. Backlash to this progressive housing program was swift. Residents in the Northern California coastal city of Eureka spearheaded a statewide constitutional amendment which would allow voters to stop their local housing authority from developing low-income housing in their community (Cavin 2019; Varian 2022).

Seeking a megaphone, the Eureka movement partnered with the California Real Estate Association (CREA), a precursor to the modern California Association of Realtors and the largest real estate group in the country. The CREA paid for the campaign to pass the measure, pitching the amendment as essential to countering "minority pressure groups," preserving white neighborhoods and therefore home values (Staff 2019). In contrast, the "pro" argument in the 1950 official voter guide did not explicitly oppose affordable housing, but instead elevated local democracy:

> A "Yes" vote for this proposed constitutional amendment is a vote neither for nor against public housing. It is a vote for the future right to say "yes" or "no" when the community considers a public housing project... It is an expression of confidence in the community's future and in the democratic process of government (*Voter Information Guide for 1950, General Election* 1950).

Passing with a narrow majority, Article 34 immediately began throttling the state's supply of new affordable housing. By 1968, voters had turned down nearly half of the public housing that had been proposed — around 15,000 units (Cavin 2019). Many public housing agencies shelved projects rather than put them to a vote. Others attempted to ameliorate voters' concerns about aesthetics and concentrated poverty. In 1968, the San Jose Housing Authority put forward a referendum in support of small duplexes and apartments of no more than four units scattered throughout the city (Cavin 2019). Still, the measure failed.

Meanwhile, the democratic appeal of Article 34 followed it beyond the ballot box. The amendment eventually arrived before the U.S. Supreme Court, where it was upheld in a 5-3 vote. Writing for the majority, the generally liberal Justice Hugo Black emphasized the democratic nature of the law, finding no evidence that the law was racially motivated: "This procedure ensures that all the people of a community will have a voice in a decision which may lead to large expenditures of local revenues. It gives them a voice in decisions that will affect the future development of their own community" (*James v. Valtierra*, 402 U.S. 137 (1971)). At the justices' private conference, Chief Justice Warren Burger allegedly scoffed at the plaintiff's claims, framing their argument as a suggestion that "too much democracy violates the Equal Protection Clause" (Cavin 2019).

Having lost at the Supreme Court, efforts to repeal Article 34 returned to the ballot box. In 1980, the pro-Article 34 campaign focused on taxes and local democracy and again defeated the repeal effort. Facing a third repeal attempt in 1993, public commentary on the 1993 repeal effort connected Article 34 to the protection of community character and quality of life. In an interview with the *Los Angeles Times*, State Assembly Member Gebe Ferguson framed support for voters' veto power over housing as both unrelated to race and shared by a diverse constituency (Martinez 1993):

> I don't think anyone can blame the general public for feeling that way. You have a drive-by shooting in Mission Viejo and then you tell [residents] you want to move in low-income people?...It's not a matter of race or income either, because low-income black communities don't want low-income housing built in their communities either, because of past experiences with that.

While Assembly Member Ferguson may have extrapolated when describing the preferences of "low-income black communities," affordable housing has been historically stigmatized due to stereotypes about its occupants (Tighe 2010). If voters view direct democracy over housing as a way to exclude poor minority residents, then the veto power granted by Article 34 would be a valuable backstop against not just affordable housing, but local racial demographic change.

Racial Demographic Change and Exclusionary Behavior

What shapes voter support for direct democracy? From a principle-based perspective, voters may believe that some decisions should be voted on directly, others should not. In this case, support for direct democracy should remain stable or change in a way orthogonal to the expected policy outcome itself. But voters are often only "weakly principled," caring more about policy outcomes than the process itself (e.g., Bartels and Johnston 2020; Prothro and Grigg 1960). From a policy-based perspective, voters may be more likely to protect direct democracy if they believe that doing so would lead to the policy outcomes they prefer, compared to turning control over to an administrative or legislative body.

Building on a policy-based perspective, voter support for direct democracy may be based on their self-identified ingroup's status in comparison to relevant outgroups. For example, a primary threat to the majority group's power is demographic change, wherein a new population threatens the status quo power structure (Blalock 1967; Blumer 1958). Threat-induced support may be less a concern about the particular policies pursued by the outgroup, but rather a general anxiety over the loss of economic, cultural, and political power (e.g., Abrajano and Hajnal 2015; Thompson 2023). Absent specific conditions found to support positive inter-group contact and limit racial bias (Allport 1954), a sudden increase in the local non-white population is expected to elicit a desire among white voters to keep direct democracy as a means to ensure exclusionary policy.

To be clear, racial demographic change may occur at the national, regional, or local level. Each level may evoke an exclusionary response through a different mechanism. National demographic change may trigger feelings about identity and culture, as well as political power over federal policy. Regional change may drive media attention and stories about crime within one's state or the nearest center city. Local demographic change may elicit greater personal self-interest via expected changes in one's home value and neighborhood trajectory. While change at multiple levels may occur simultaneously, local demographic change should most closely drive exclusionary behavior via control over local housing policy as in the case of Article 34.

Demographic Change in Los Angeles County, 1993

The November 1993 vote to repeal Article 34 is an opportunity to assess the importance of local racial demographic change across a racially diverse array of communities. From 1980 to 1993, Los Angeles County — the focus of this study — saw relatively little growth in its Black population. Instead, the county experienced a large increase in Latino migration and a proportionally large increase in Asian migration mainly from outside of the United States. Simultaneously, Los Angeles County's non-Hispanic white population decreased from 53 to 38 percent of the population. From an individual's perspective, the median Los Angeles County voter experienced an 14 percentage point (54 percent) increase in the non-white population share of their precinct.

For local demographic change to trigger racial threat, voters must perceive their neighborhoods as changing (Hopkins 2009; Wong 2007). A 1992 survey of Los Angeles County residents found that 56 percent of respondents felt that their neighborhood was experiencing a change in ethnic composition (Bobo and Zubrinsky 1996). Furthermore, this influx of Latino and Asian residents was concentrated in areas which had been majority Black going back to the 1950s (Bergesen and Herman 1998). As a result, Black residents were the racial group most likely (78 percent) to agree that their neighborhood was racially changing (Bobo and Zubrinsky 1996).

This local racial demographic change may lead to exclusionary behaviors through two main pathways. First, while members of many racial groups tend to show a degree of in-group preference (Charles 2006; Farley et al. 1994; Krysan et al. 2009), there may also be a preference for having a sizable share of white neighbors due to concerns for their neighborhood's political status and socioeconomic trajectory. White constituents often receive better representation, responsiveness, and public goods provision from local elected officials (e.g., Hankinson and Magazinnik 2023; Schaffner, Rhodes and La Raja 2020). At the same time, residents' opposition to demographic change may also come from who moves in as white residents leave. Middle-class minority communities may struggle to preserve socioeconomic status, especially in segregated contexts where they are spatially proximate to poorer, majority-minority neighobrhoods. Pattillo-McCoy (1999, p. 6) writes:

> Middle-income black families fill the residential gap between the neighborhoods that house middle-class whites and the neighborhoods where poor African Americans live. Unlike most whites, middle-class black families must contend with the crime, dilapidated housing, and social disorder in the deteriorating poor neighborhoods that continue to grow in their direction. Residents attempt to fortify their neighborhoods against this encroachment...

Thus, turnover from white to non-white neighbors may signal a risk of declining political influence, city services, and even home values.

As evidence of these concerns in the context of Los Angeles, Charles (2006)'s contemporaneous research used an experimental design wherein respondents were asked their willingness to move into stylized neighborhoods of varying racial composition. Known as the Los Angeles Study of Urban Inequality (LASUI), the survey was fielded from 1992 to 1993, making its conclusions especially relevant for understanding the November 1993 election in Los Angeles County. Documenting these preferences, Charles (2006, p. 183) writes:

> Across racial groups, patterns of neighborhood racial composition preferences reveal a clear and consistent racial rank-ordering of out-groups as potential neighbors. Whites are always the most preferred out-group neighbors..."

Charles finds that this preference hierarchy even extends to neighborhood change among non-white residents: "Across racial groups, blacks are indisputably the least-desired neighbors. It is equally clear that, among non-whites, integration with whites is more favorable than integration with other non-whites" (Bobo et al. 2000, p. 193).

Second, voters may be as responsive to changing local demographics due to perceived competition among racial groups for space and resources (Bobo and Hutchings 1996; Gay 2006). Drawing from her research in Los Angeles County, Charles (2006, p. 164) notes: "Concerns about relative group position are also somewhat apparent among blacks, whose preferences for both Asian and same-race neighbors are negatively influenced by the belief that this group poses a competitive threat to economic opportunities and political power." Contemporary reporting reflects this conflict. A *Los Angeles Times* article written months before the 1992 Rodney King riots captures the unease among Black, Latino, and Asian residents. Clifford (1991) writes:

> Cultural collisions, often violent, occasionally fatal, are occurring every day. Hostilities between black residents and Korean shop-keepers, Latinos and blacks vying for jobs at Martin Luther King Jr./Drew Medical Center, interracial fighting at Lawndale high school, and repeated charges of police brutality against minorities — all of this is disturbing the city's racial peace in a way that has some political analysts recalling Watts.

Additional reporting at the time highlights the competition felt by long-time Black residents towards the new arrivals. Noble (1995) notes:

> Marilyn Thompson, 45, a [Black woman and] telephone company manager who lived in Los Angeles for nearly 20 years before moving to Atlanta last year, also said the doors of opportunity seemed open for Hispanic and Asian residents but shut for blacks in California. "It seems like you can come into California and have nothing and end up with everything," she said of other immigrants to the state. But, she said, blacks "can't seem to get ahead."

How does this perceived racial hierarchy and competition over housing and jobs shape intergroup relations? Eighteen months prior to the 1993 election, four white Los Angeles police officers were acquitted in the beating of Rodney King, a Black man. Rioting began within hours, concentrated in South Central Los Angeles and Koreatown. Researchers argue that the riots highlighted competition-fueled resentment between racial groups, with Black and Latino rioters targeting the Korean community. Black residents viewed Latinos as having taken over their community and competing in the labor market, whereas Latinos believed they were underrepresented politically (Johnson Jr and Farrell Jr 1992). Bergesen and Herman (1998) attribute the intensity of the riots to this desegregation and neighborhood succession, with areas of racial demographic change showing the highest rates of violence.

In short, much of the existing literature associates racial demographic change with feelings of racial threat among white voters. However, local demographic change can also provoke exclusionary behaviors among non-white voters. Shared concerns may be grounded in what a shrinking white population means for a community's political representation, socioeconomic status, and access to public goods. Or, the response can be directed towards the arriving group, either driving competition for resources or animating both inter-racial group tensions. Through both pathways, increases in the local non-white population may drive support for the voters' veto over affordable housing among voters broadly.

Effect of Demographic Change

To test the effect of racial demographic change on support for the voters' veto, I use two analytical strategies. First, I assess the effect of within-precinct demographic change from 1980 to 1993 on a precinct's support for repealing Article 34 in the 1993 November election. Second, I subset to precincts which experienced minimal residential churn and were racially stable from 1980 to 1993, then measure the effect of spatial proximity to the nearest rapidly diversifying Census tract on support for repealing the voters' veto. Results from both strategies are substantively the same. With these two approaches, I also show that the effect of demographic change on exclusionary behaviors is shared among both white and non-white voters. I begin with the within-precinct design.

Data

I combine precinct-level election returns with tract-level Census data to generate 3,718 precinct-level observations in Los Angeles County. Compared to citywide returns, precinct-level measurement provides substantially more statistical power and is more likely to accurately reflect the variation in residents' local experiences. The analysis is limited to Los Angeles County as that is the only county where I have been able to find the contemporaneous voter file required for the analysis. Still, Los Angeles County provides a uniquely dense multiracial context which allows me to observe the effects of demographic change on a diverse array of communities. Also, as of 1993, Los Angeles County contained 29% of California's population, making this a substantively meaningful subset of the state's voters.

Combining precinct returns with Census data presented a challenge. Los Angeles County does not have a shapefile of 1993 precinct boundaries. Thus, I use Enos, Kaufman and Sands (2019)'s geocoded Los Angeles County voter file from 1992 which includes each voter's address and precinct. I overlay the voters on shapefiles of the 1980, 1990, and 2000 Census tracts (Manson et al. 2022), generating data for the total number of voters per tract and per tract-precinct sub-unit. These quantities allow me to estimate the share of each tract's population that can be attributed to each precinct. I use this ratio to allocate counts of other Census data, e.g., the number of non-white residents, the number of manufacturing employees, etc. After allocating these tract-level counts to each tract-precinct sub-unit, I sum the counts within each precinct. Using these precinct-level counts, I then calculate the percentages and values needed for the analysis.

Dependent Variable My dependent variable is based on the precinct-level vote share for the repeal of Article 34 as voted on in November 1993. Within the precinct data, the weighted mean vote share was 44.6% in favor of repeal. For conceptual clarity, I define *support for the voters' veto over housing* as the complement of support for repealing Article 34. Rather than contextualize effects as one's opposition to repealing the amendment, I can say that 55.4% of voters supported direct democratic control over new affordable housing proposed for their community.

Figure B-1 shows the distribution of support for this voters' veto across Los Angeles County. Blue points are the centroids of precincts that are in the top tercile of support ($\geq$62% in favor of the voters' veto), orange points are precincts in the bottom tercile of support ($\leq$47% in favor), and the middle tercile is omitted for visualization purposes only. Support for the voters' veto is concentrated outside of the central, more urbanized areas of Los Angeles County. The spatial trend largely aligns with partisanship as found using voter file party registration. Precincts with higher rates of Republican registered voters were also more likely to support the voters' veto (Figure B-2).

Independent Variables The treatment is local racial demographic change in the lead-up to the 1993 election. Given the election occurred prior to the creation of the American Community Survey in 2005, the data are limited to the decennial Census. To estimate values for 1993, I linearly interpolate data from the 1990 and 2000 Censuses. I operationalize this change in two ways:

- Percentage point change ($Pop_t - Pop_{(t-1)}$)
- Proportion change ($\frac{Pop_t - Pop_{(t-1)}}{Pop_{(t-1)}}$)

For reference, the population-weighted median precinct experienced an 14 percentage point (54 percent) increase in non-white population from 1980 to 1993. It is theoretically unclear which definition better captures local demographic change, so results are reported for both (e.g., Hill, Hopkins and Huber 2019). However, the correlation between the two measures of demographic change is 0.65, suggesting that the treatments are capturing different aspects of a similar phenomenon. Section C discusses in detail the ways in which these two measures differ in the context of Los Angeles County.

Given this treatment is estimated by aggregating tracts into precincts, there is inevitable measurement error in operationalizing demographic change from 1980 to 1993. Even more, because the treatment is derived from two precinct-level estimates (1980 and 1993), the two opportunities for measurement error may compound each other, producing outlier swings in demographic change which do not reflect reality. While these errors are not expected to introduce bias, I account for them by presenting all results using winsorized treatments — truncating extreme values at both ends to their 5th and 95th percentile, respectively. The results using non-winsorized data are substantively the same, statistically significant, and reported in Table L-18.

Empirical Strategy

I regress precinct-level vote share in support of the voters' veto on the 1980-1993 change in non-white population at the precinct level. For comparability, I operationalize both treatments as a standard deviation increase in the precinct's non-white population. The identifying assumption of this strategy is selection on observables. Thus, I selected theoretically motivated control variables to account for other changes which may confound the relationship between an increase in the local non-white population and increasing support for the voters' veto over affordable housing.

A spurious relationship between the two may occur if the areas where voters are concerned about social and economic conditions in 1993 were also more likely to attract non-white residents from 1980 to 1993. For example, economic traits may predict both economic anxiety and an influx of immigration. Thus, I control for the precinct-level unemployment rate, vacancy rate, and household median income in 1980. Other observable traits may be associated with residents' ability to access local government resources and organize to stabilize their neighborhood, such as the percent of residents who are homeowners, the local residential density, and the percent of residents in 1980 who were white. If these traits are inversely correlated with the likelihood of attracting immigration, we may see an spurious relationship.

To review, the models include controls for each precinct's 1980 pretreatment percent non-Hispanic white, homeownership rate, vacancy rate, log median household income, population density, and unemployment rate. Tables K-16 and K-17 show regression results using both the 1980 pretreatment level of each control as well as the change in each control from 1980 to 1993. While including changes in controls over time biases the effect of demographic change as a treatment, results are substantively identical and statistically significant.[4]

I also condition for time-invariant place attributes by using a Census place-level fixed effect and weight the data by precinct population. As shown in Figures C-4 and C-5, local racial demographic change across precincts is non-random. Non-white residents may be more likely to select Census places to live, rather than selecting specific precincts. To account for this correlation across observations, I cluster Huber-White standard errors at the Census place-level.

Results

Table 1 presents the effect of a standard deviation increase in the non-white population on support for the voters' veto over affordable housing.[5] Results for a percentage point change are shown in Models 1 and 2, while results from a proportion change are in Models 3 and 4. Models 1 and 3 use only Census place-level fixed effects, whereas Models 2 and 4 add controls for pretreatment levels. A standard deviation increase in the percent non-white population (+10.4 percentage points) is associated with a 4.5 point (d = .32) increase in support for direct democracy over housing. A standard deviation increase in the non-white population as a proportion change (+54.2 percent) is associated with a 3.4 point (d = .24) increase in support for the voters' veto. These effects are nearly identical to those using non-winsorized

[4] I do not include a control for precinct-level partisanship, as I do not have pretreatment levels of partisanship.

[5] Table D-1 presents the Table 1 including the coefficients for each control variable. Figure I-9 visualizes the bivariate relationship between the two treatment variables and precinct support for the voters' veto using the raw data.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.054***	0.045***		
	(0.014)	(0.011)		
Prop. Δ in non-white pop.			0.069***	0.034***
			(0.008)	(0.010)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.573	0.751	0.613	0.723
Adj. R^2	0.556	0.740	0.598	0.711
Num. obs.	3718	3651	3718	3651
RMSE	3.333	2.556	3.173	2.694
N Clusters	144	142	144	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table 1: Effect of change in a precinct's non-white population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized.

data and are stable across an array of model specifications, such as the removal of controls and place-level fixed effects (see Tables K-16 and K-17).

While I cluster standard errors at the place-level to account for correlated treatment within places, there may be other omitted variables that contribute to the error term and are spatially correlated. Likewise, this study's proximity-based analysis (below) is premised on the belief that precinct outcomes are affected by nearby Census tracts. Both conditions violate SUTVA. As a robustness check, I use a spatial Durbin error model (SDEM) to account for potential spillover effects from neighboring precincts which may be affecting the relationship, as well as spatial autocorrelation among standard errors. The results are statistically significant and substantively similar as the non-spatial OLS models (Section H).

As further evidence of the importance of racial demographic change, I replicate Table 1 with variables capturing alternative explanations. Tables J-12 and J-13 show that support for the voters' veto is related to neither precinct-level changes in population density nor residential churn, defined as the share of the population that moved to their unit in the past 10 years. While Table J-14 shows that a decrease in the local household median income is associated with an increase in support for the voters' veto, the magnitude of the effects is roughly half that of those from racial demographic change. Lastly, much of California's increasing diversity in the 1980s came from immigration, with undocumented immigrants alone representing 22 to 31 percent of all migrants to CA during this period (Johnson 1996). However, Table J-15 does not detect a consistent treatment effect from an increase in the percent foreign born within a precinct.

Mechanisms by Racial Groups

As discussed, much of the literature on racial demographic change has emphasized the sensitivity of white voters to an increase in the local non-white population. Accordingly, I had expected that the largest effect of demographic change would be found in the precincts with the largest white population shares as of 1980. To assess this differential response, I divide precincts into terciles by their pretreatment white population. I then replicate the models in Table 1, but interact treatment with an indicator for whether a precinct falls in the top tercile of percent white ($\geq$83%). I drop the middle tercile of pretreatment white population. The bottom tercile includes precincts which were less than 55% white as of 1980. I also omit the control for pretreatment percent white, due to collinearity with the percent white tercile indicator.

Across both definitions of demographic change, a precinct's white population in 1980 is a significant moderator of the treatment, but not in the way hypothesized (Table 2).[6]. The effect of demographic change is either equal to or substantively and statistically *larger* in more racially diverse precincts (the lower order treatment term). In contrast, "white precincts" show much more muted effects.

To better understand this relationship among specific racial groups, I disaggregated precincts by their pretreatment racial composition. Majority Latino and majority Black precincts express a similar increase in support for the voters' veto in response to an increase in their local non-white population, as compared to non-white precincts broadly (Figure E-

[6]Table D-2 presents the Table 2 including the coefficients for each control variable.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.049***	0.048***		
	(0.007)	(0.007)		
Prop. Δ in non-white pop.			0.134***	0.123***
			(0.024)	(0.021)
Pct. point Δ x white precinct, 1980	−0.021*	−0.009		
	(0.008)	(0.008)		
Prop. Δ x white precinct, 1980			−0.112***	−0.094***
			(0.017)	(0.015)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.668	0.757	0.663	0.747
Adj. R^2	0.649	0.743	0.644	0.732
Num. obs.	2454	2418	2454	2418
RMSE	3.103	2.663	3.126	2.718
N Clusters	129	127	129	127

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table 2: Effect of change in non-white population (1980-1993) on support for voters' veto, interacted with percent white at precinct level, 1980. Treatment data are winsorized.

8).[7] I also use King (1997)'s Ecological Inference method to estimate the effect of demographic change on behavior by racial group. As shown in Section F, I find a positive effect of demographic change on support for the voters' veto among both white and non-white voters.

Finally, to show more direct evidence that non-white residents express an exclusionary response to local racial demographic change, I geocoded three recent surveys containing over 8,000 respondents. Each survey contained a question measuring the outcome of interest: support for local control over housing. As discussed in Section G, I replicate the approach of the observational data by regressing support for the local control on the increase in ZIP code-level racial demographic change over the preceding 14-year period (2007 to 2020). I use a set of controls which were common across the three surveys, including homeownership, income, education, and ZIP code percent white pretreatment (2007). I also include a survey-level fixed effect and use Huber-White standard errors.

[7] Given there were only 8 majority Asian precincts in 1980, an independent analysis of majority Asian precincts is not feasible.

These individual survey results match the pattern of the observational data. Respondents in ZIP codes which experienced a recent increase in their non-white population show greater support for local control over housing (Table G-9). This support is either shared across residents or driven by non-white residents. To limit the risk that respondents are recent arrivals, I replicate the analysis using only homeowners, who are more likely to be residentially stable, have experienced a longer duration of this demographic change, and less able to select out of their location as it is changing. Results from homeowners replicate the pattern from the full sample (Table G-10). The link between local demographic change and exclusionary attitudes is as strong if not stronger among non-white respondents compared to white respondents.

Of note, voters may respond differently to influxes of specific racial groups over this period. But disaggregation of the treatment into individual racial groups has empirical limitations. A precinct that increases in the Latino population may also be increasing the Asian population at the same time. Isolating only a share of this overall treatment will generate a biased estimate. Still, as discussed in Section E, it appears that voters respond to increases in Latino, Black, and Asian residents, suggesting the importance of a decrease in the local share of the white population rather than an influx of any specific racial group.

Finally, I theorized that this exclusionary response may be driven by one of two mechanism: a concern over a shrinking white population or a decrease in one's own racial group. To evaluate these theories, I test whether majority Black and Latino precincts show a similar increase in support for the voters' veto when their own racial group decreases in size. As shown in Tables E-5 and E-6, a decrease in these minority groups' own population share does not elicit an exclusionary reaction. Only a decrease in the local white population spurs support for local control over affordable housing. These subgroup analyses suggest that both white and non-white voters may be growing concerned over the economic, social, and political trajectory of their neighborhoods as the local white population shrinks, rather than over the size and influence of their own racial group.

Effect of Proximity to Demographic Change

One threat to the above analysis is the possibility that voters within a diversifying precinct were not changing their behavior, but rather the composition of the precinct's electorate was changing. In other words, the higher support for the voters' veto may be coming from new voters arriving in these precincts.

This alternative explanation is unlikely for two reasons. First, precincts with more non-white voters on average show lower support for the voters' veto (Figure B-3). Thus, it is improbable that the newly arrived non-white residents would be *more* likely to vote in favor of the voters' veto compared to long-time white residents. Second, for residential churn to increase support for the voters' veto, those leaving the precinct would have to be *less* likely to support direct democracy over housing than those staying behind. This is also improbable. Exit in response to local demographic change is more likely among the least racially tolerant (Clark 1991) — those who are the most supportive of the voters' veto. If anything, I would expect to see less support for the voters' veto in diversifying areas if the least racially tolerant were being replaced with new arrivals.

Another way to account for the possibility that these effects are driven solely by residential turnover is to use an identification strategy which does not rely on racial change *within* each precinct as the treatment. Instead, I measure exposure to demographic change via the proximity of each precinct to a rapidly diversifying Census tract (e.g., Reny and Newman 2018). Across a variety of specifications, I find that racially stable, low residential churn precincts closer to diversifying neighborhoods were more supportive of the voters' veto over affordable housing compared to those farther away. This findings suggests that residential turnover is unlikely to be responsible for the results observed in the previous section.

Data

Using the same precinct data as described above, I calculate each precinct's proximity to a "diversifying tract," a Census tract which experienced an extreme increase in its non-white population share.

Dependent Variable The dependent variable remains precinct-level vote share in support of the voters' veto over affordable housing in 1993.

Independent Variable The treatment is the proximity of each precinct to a Census tract which experienced an extreme increase in non-white population. For my main specification, I define diversifying tracts as those exceeding the 90th percentile of the increase in the non-white population from 1980 to 1993. For the percentage point treatment, this is a >27 percentage point increase in non-white population. For the proportion change, this a >127 percent increase in the non-white population. Results are substantively similar using cutpoints at the 85th and 95th percentiles (See Tables M-19 to M-22). I then measure proximity as the distance between the centroid of each precinct and its nearest diversifying tract.

One challenge of using Census tracts to identify diversifying areas is that they do not perfectly match known neighborhood boundaries, meaning voters' perceptions of change may not match the administrative boundaries of Census tracts (Wong 2007). On the other hand, Census demographers are careful to construct tract boundaries based on known landmarks, such as rivers, rail lines, and major streets. These boundaries tend to be very stable over time and are the most widely used unit of observation in quantitative studies of neighborhood racial change (Lee et al. 2008).

Still, some tracts have few residents, meaning a large percent change in the non-white population may be neither substantively meaningful nor perceptible to those nearby. After defining the percentile cutpoints for demographic change, I subset to tracts outside of the bottom tercile in population (>3,136 residents).[8] As a result, the set of diversifying tracts are those which experienced large increases in their non-white population share from 1980-1993 and were also large enough that their demographic change should register in voters' minds. Figure C-7 shows the distribution of these tracts using both the percentage point change and proportion change definitions.

Empirical Strategy

In this design, the treatment is a precinct's proximity to a diversifying Census tract — one which diversified significantly from 1980 to 1993. As in the previous models, I control for the same pretreatment (1980) precinct-level covariates that would confound the relationship, including percent non-Hispanic white, homeownership rate, log median household income, vacancy rate, unemployment rate, and population density. Because treatment is based on proximity to a diversifying tract, I cluster Huber-White standard errors at the level of the nearest diversifying tract.

Results

I transform "distance away from" into the more analytically useful "proximity to" by multiplying distance by -1. I then standardize proximity so that the treatment is a standard deviation increase in proximity to a diversifying Census tract, centered at the mean. For context, a standard deviation increase in proximity is approximately 3.5 kilometers (2.2 miles). I regress support for the voters' veto in 1993 on proximity to a diversifying tract, weighting by population.

The purpose of this analysis is to test whether precincts which experienced little change in their demographics from 1980 to 1993 were responsive to demographic changes around them. If so, that would suggest that the within-precinct effects are unlikely to be driven by residential turnover. To test this, I subset to precincts which are in the bottom tercile of

[8]Results are substantively similar when relaxing this threshold to the bottom quartile.

both residential churn and change in their own non-white population from 1980 to 1993.

	Percentage Point Δ		Proportion Δ	
	Model 1	Model 2	Model 3	Model 4
Prox. to pct. point Δ in non-white	0.030***		0.018**	
	(0.008)		(0.006)	
Prox. to prop. Δ in non-white		0.052***		0.040***
		(0.008)		(0.007)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes
R^2	0.597	0.665	0.595	0.643
Adj. R^2	0.552	0.628	0.545	0.598
Num. obs.	403	403	373	373
RMSE	2.333	2.125	2.434	2.286
N Clusters	37	37	33	33

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table 3: Effect of proximity to diversifying tract on support for voter's veto, among precincts which changed minimally over the past ten years.

Table 3 shows the effect of proximity to a diversifying tract among precincts in the bottom tercile of demographic change.[9] For Models 1 and 2, these are precincts where the non-white population increased by less than 6.5 percentage points (median change of +3 percentage points non-white). For Models 3 and 4, the non-white population increased by less than 27 percent (median change of +4 percent non-white). I again vary whether the nearby diversifying tracts are defined based on percentage point (Models 1 and 3) or proportion changes (Models 2 and 4). All models include the same pretreatment covariates as well as Census place fixed effects.

Within precincts which saw little demographic change themselves, areas which were closer to diversifying Census tracts were on average 3.5 points (d = .24) more supportive of the voters' veto over housing compared to those farther away. These effects do not vary between predominantly white and non-white precincts, again demonstrating a broadly shared reaction to local racial demographic change (Table N-23). Unfortunately, the additional subsetting by demographic stability means that independent analyses on majority Latino, majority Black,

[9]Table D-3 presents the Table 3 including the coefficients for each control variable.

or majority Asian precincts is not feasible with this design.

Discussion

These findings are timely. Los Angeles County of 1993 was a bellwether for California as a whole today. In the 1990 Census, Los Angeles County was 41% non-Hispanic white, 11% non-Hispanic Black, and 10% Asian; 38% of residents identified as Latino. As of 2020, California as a state is nearly identical: 41% non-Hispanic white, 6% non-Hispanic Black, and 15% Asian; 39% of Californians identify as Latino. The housing costs and competition that Charles (2006) documented in Los Angeles have now spread across the state, with California having the highest housing costs in the nation.

Yet despite these high costs, California continues to diversify. From 2010 to 2020, California's white population decreased by 24 percent, the largest statewide decrease. Coincidentally, from 1980 to 1990, Los Angeles County's white population also decreased by 23 percent. These parallel paths may inform politics today. In November 2024, Californians were scheduled to vote again on repealing Article 34. However, the state's Democratic Party leaders pulled the ballot measure, under the guise of too many competing initiatives on the upcoming ballot. In reality, polling found the measure underwater in terms of public support (Board 2024). Given the similarities in demographic trends leading up to 1993 and 2024, advocates may have been wise to pull the measure and avoid a rebuke of the state's efforts to limit control over housing.

Strategic awareness around state preemption and limiting the voters' veto is important. Be it via ballot box or legislature, local control of land use has been a tool of segregation. Hopes to loosen local control to address both racial segregation and the supply-side of the housing affordability crisis hang on not only weakening the voters' veto, but on scaling up control to the state level. But these findings should give us pause: increasing local racial diversity begets greater support for local democracy in housing. And while the CA state government struggles to repeal Article 34, a countermovement — "Our Neighborhood Voices" — is organizing for a competing initiative that would effectively end state preemption of local land use policy (Brasuell 2022). With California continuing to experience demographic change, voters may seize the opportunity to solidify local control over housing and thus integration.

Conclusion

As Western democracies diversify, our understanding of the political consequences of racial demographic change has been limited in three ways. First, research has focused heavily on the behaviors and attitudes of the majority population, emphasizing racial threat as the dominant mechanism behind observed relationships. Second, our understanding of exclusionary reactions to racial change has been most commonly observed via survey attitudes or indirect behavioral outcomes, such as voter turnout. Third, researchers have emphasized support for national policy outcomes like restrictive immigration measures, rather than the exclusionary tools used at local context, where voters may be most directly exposed to demographic change.

This study is a direct response to these three shortcomings. The attempted repeal of Article 34 is a rare case of observing voters' behavioral support for exclusionary policy, in this case the voters' veto power over the historically racialized policy of affordable housing. Using precinct-level returns and multiple analytical approaches, I show that local exposure to an increasing non-white population leads to greater support for voter control over affordable housing. For these voters, direct democracy over affordable housing appears to be a means of defense; "fortifying" their communities in the face of potential political, economic, and social changes.

Furthermore, this exclusionary behavior was not only expressed by majority white precincts, but across precincts with already sizeable Black, Latino, and multiracial populations. That this impulse was shared across an array of racial contexts suggests that understanding the effects of growing diversity in Western democracies requires more attention to theory beyond white racial threat. In the case of Los Angeles County, contemporary research suggests that broad swaths of voters may have associated the shrinking local white population with concerns for their neighborhood's social, economic, and political trajectory. My findings support this mechanism, as majority Black and Latino precincts did not show similar exclusionary behaviors in response to a decline in their own group's share of the precinct population.

Currently, Americans harboring exclusionary attitudes — either on national immigration policy or local zoning practices — are more likely to find their policy preferences embraced by the Republican Party. As the United States diversifies, there is growing evidence that many Latino, Asian, and to a lesser extent Black voters may be to moving away from the Democratic Party, especially in areas with high levels of immigration. The exclusionary behaviors found in this study align with and may provide a local foundation for this national shift of non-white voters towards the Republican Party. More broadly, these findings and California's struggle to repeal the voters' veto underscore the tension between integration and local democratic control over housing in a diversifying society.

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Online Appendix for "The Voters' Veto: Local Racial Demographic Change and Exclusionary Behavior"

Contents

A Article 34 Background

Article 34 of the California Constitution:

> "No low rent housing project shall hereafter be developed, constructed, or acquired in any manner by any state public body until, a majority of the qualified electors of the city, town or county, as the case may be, in which it is proposed to develop, construct, or acquire the same, voting upon such issue, approve such project by voting in favor thereof at an election to be held for that purpose, or at any general or special election."

1993 ballot measure to repeal Article 34 (Proposition 168):

> "LOW RENT HOUSING PROJECTS. LEGISLATIVE CONSTITUTIONAL AMENDMENT. Amends state constitutional definition of low rent housing projects to include only projects owned by a governmental entity as defined. Excludes projects found to have no significant negative impact on the revenues of the affected governmental entity, and whose physical appearance is found to have no significant negative impact on the surrounding community. Requires approval by voters only upon qualification of ballot petition as specified. Exempts projects approved on or before November 3, 1992, or projects with existing contracts for federal financial assistance."

B Descriptive Relationships

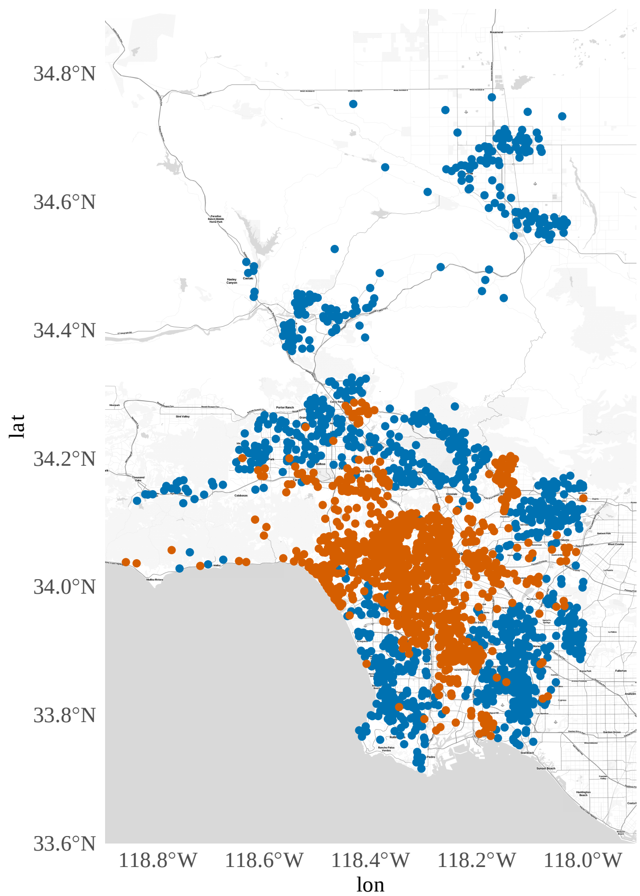


Figure B-1: Distribution of support for the voters' veto using precinct centroids. Blue as the top tercile of support (≥62% in favor); orange as the bottom tercile of support (≤47% in favor). The middle tercile is omitted for visualization purposes only.

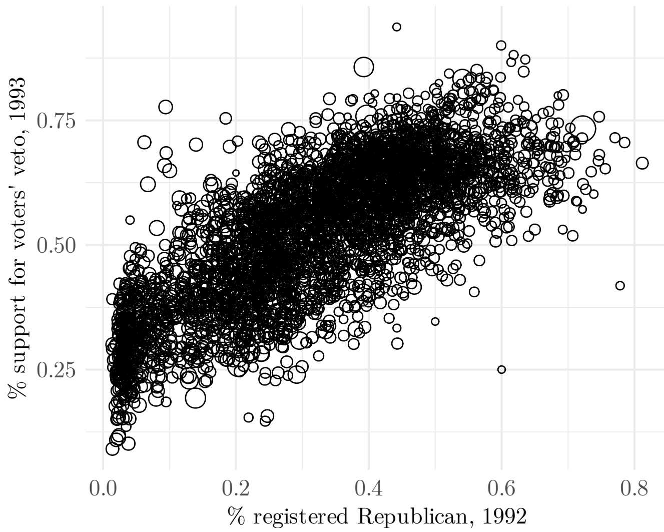


Figure B-2: Relationship between precinct-level Republican registration in 1992 and support for Article 34 repeal in 1993.

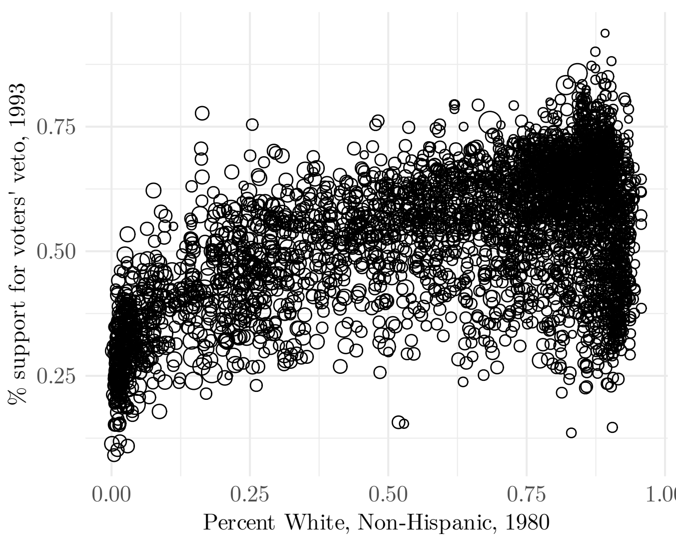


Figure B-3: Relationship between precinct-level percent white, non-Hispanic (1980) and support for Article 34 repeal (1993).

C Defining Treatment

Figures C-4 and C-5 show maps of how precincts changed from 1980 to 1993. Green precincts experienced the smallest increase in their non-white population (bottom tercile), pink precincts experienced the greatest increase in their non-white population (top tercile), with the middle tercile omitted. Q3 precincts (pink) of both measures of racial change generally cover comparable areas, though with some notable exceptions around the Santa Monica Mountains in the upper left quadrant of each map.

To understand how areas of intense demographic change may differ, Figure C-6 shows the relationship between the 1980 percent non-Hispanic white population (i.e., pretreatment white population) on the x-axis and each treatment measure on the y-axis. In very white areas, large changes in percentage point population are unlikely, e.g., 98 percent white precincts cannot experience 20 percentage point increases in non-white population. However, a large proportion change is common in very white areas. A 98 percent white precinct could easily experience a 4 percentage point increase in the non-white population, equaling a 200 percent change in the non-white population. In short, most of the variation in the percentage point change in population occurs in areas between 50 and 80 percent white, whereas most of the variation in proportion change is in areas that are greater than 80 percent white circa 1980.

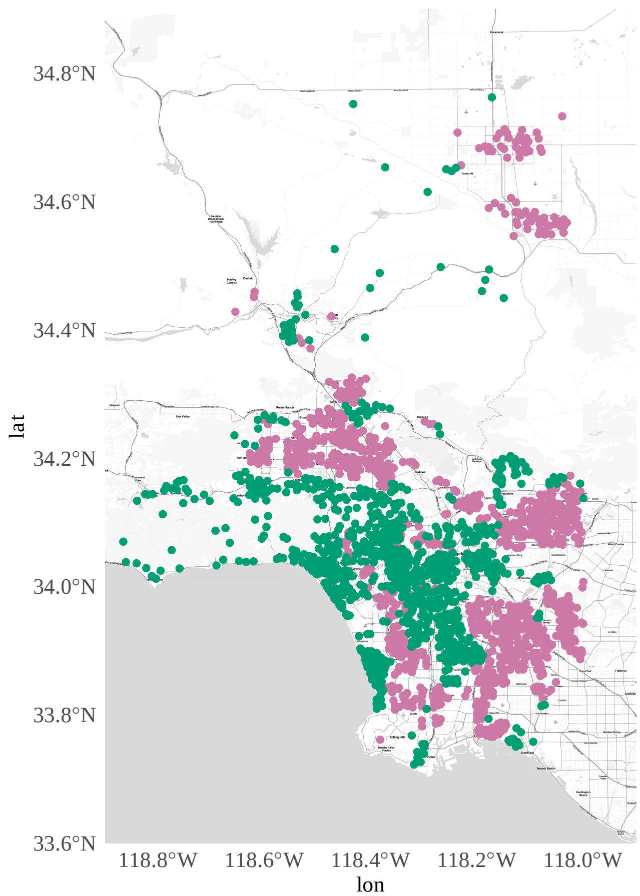


Figure C-4: Distribution of increase in non-white population as a percentage point change, 1980 to 1990 shown using precinct centroids. Green as bottom quartile of racial change; pink as top tercile of racial change. Middle tercile omitted.

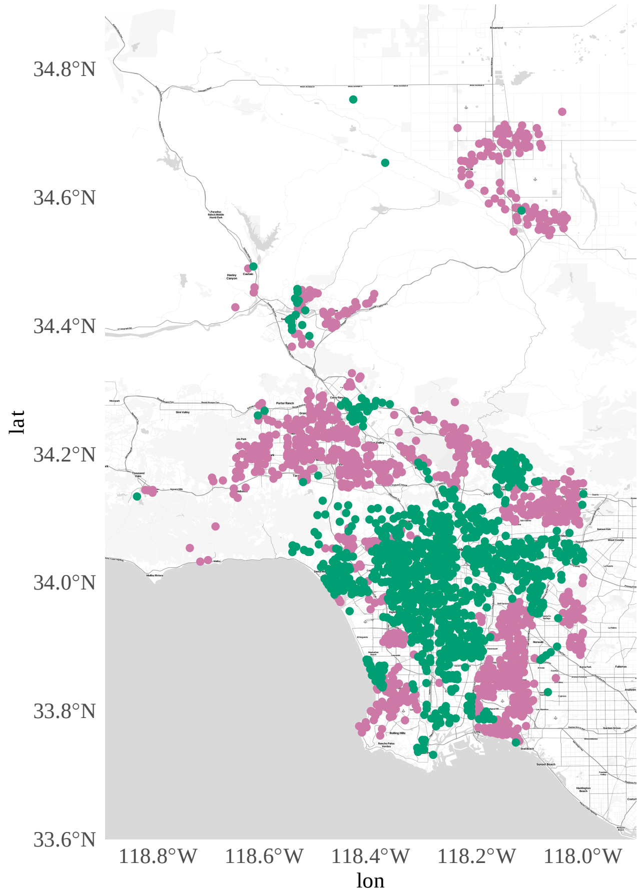
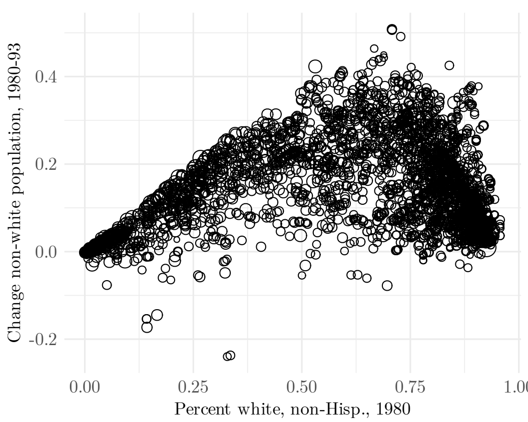
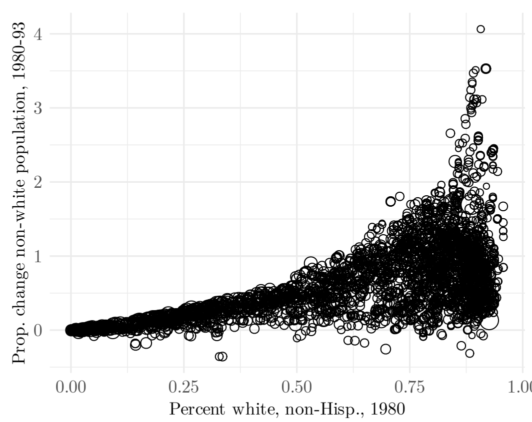


Figure C-5: Distribution of increase in non-white population as a proportion change, from 1980 to 1990 shown using precinct centroids. Green as bottom quartile of racial change; pink as top tercile of racial change. Middle tercile omitted.

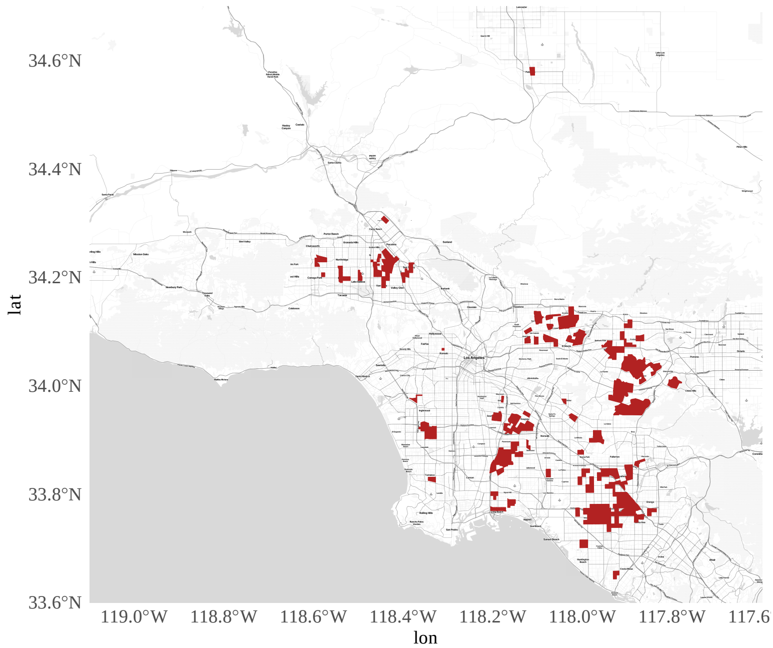


(a) Percentage point change

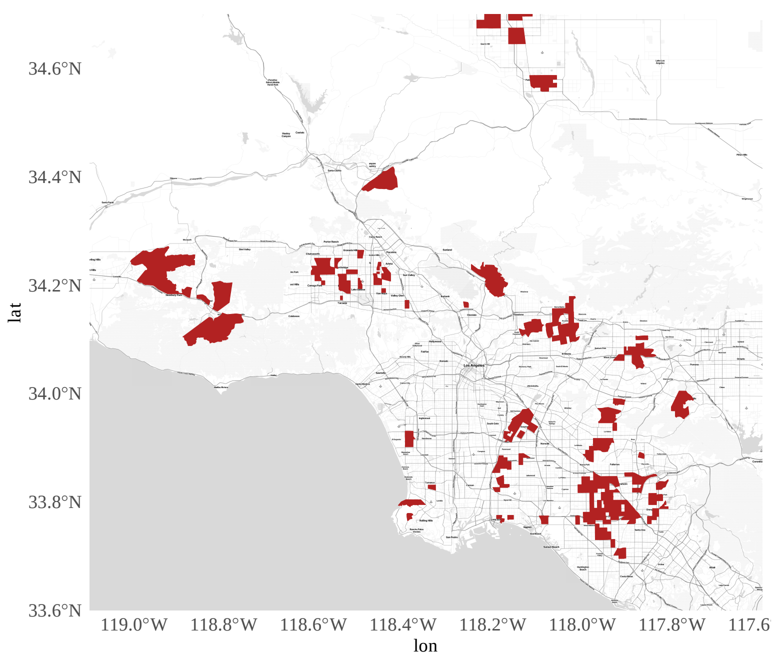


(b) Proportion change

Figure C-6: Variation in the two treatment variables based on a precinct-level percent white, non-Hispanic population in 1980 (pretreatment).



(a) Percentage point change



(b) Proportion change

Figure C-7: Census tracts above the 90th percentile for their increase in non-white population share, 1980-1993. Distance to the nearest of these tracts is used as the treatment for the for the proximity analysis.

D Main Results with Controls

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.054^{***}	0.045^{***}		
	(0.014)	(0.011)		
Prop. Δ in non-white pop.			0.069^{***}	0.034^{***}
			(0.008)	(0.010)
Pct. white, non-Hispanic (1980)		0.067		0.055
		(0.035)		(0.040)
Pct. unemployed (1980)		-0.517^{***}		-0.670^{**}
		(0.143)		(0.223)
Pct. homeowner (1980)		0.137^{***}		0.123^{**}
		(0.039)		(0.038)
Pct. vacant (1980)		-0.303^{**}		-0.274^{*}
		(0.113)		(0.113)
Log median household income (1980)		0.010		−0.004
		(0.014)		(0.015)
Pop. density (1980)		-6.040^{***}		-5.367^{***}
		(0.528)		(0.783)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.573	0.751	0.613	0.723
Adj. R^2	0.556	0.740	0.598	0.711
Num. obs.	3718	3651	3718	3651
RMSE	3.333	2.556	3.173	2.694
N Clusters	144	142	144	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table D-1: Effect of change in a precinct's non-white population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.049***	0.048***		
	(0.007)	(0.007)		
Prop. Δ in non-white pop.			0.134***	0.123***
			(0.024)	(0.021)
Pct. point Δ x white precinct , 1980	−0.021*	−0.009		
	(0.008)	(0.008)		
Prop. Δ x white precinct , 1980			−0.112***	−0.094***
			(0.017)	(0.015)
Pct. unemployed (1980)		−0.610***		−0.669***
		(0.121)		(0.141)
Pct. homeowner (1980)		0.116***		0.109**
		(0.033)		(0.035)
Pct. vacant (1980)		−0.378*		−0.370*
		(0.163)		(0.158)
Log median household income (1980)		0.020		0.023
		(0.015)		(0.015)
Pop. density (1980)		−5.814***		−4.848***
		(0.716)		(0.598)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.668	0.757	0.663	0.747
Adj. R^2	0.649	0.743	0.644	0.732
Num. obs.	2454	2418	2454	2418
RMSE	3.103	2.663	3.126	2.718
N Clusters	129	127	129	127

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table D-2: Effect of change in non-white population (1980-1993) on support for voters' veto, interacted with percent white at precinct level, 1980. Treatment data are winsorized.

	Percentage Point Δ		Proportion Δ	
	Model 1	Model 2	Model 3	Model 4
Prox. to pct point. Δ in non-white	0.030***		0.018**	
	(0.008)		(0.006)	
Prox. to prop. Δ in non-white		0.052***		0.040***
		(0.008)		(0.007)
Pct. white, non-Hispanic (1980)	0.196***	0.171***	0.149**	0.148**
	(0.049)	(0.038)	(0.047)	(0.045)
Pct. unemployed (1980)	−0.606	−1.010***	−0.799*	−1.109***
	(0.345)	(0.265)	(0.315)	(0.261)
Pct. homeowner (1980)	0.033	0.041	−0.018	−0.000
	(0.057)	(0.051)	(0.053)	(0.052)
Pct. vacant (1980)	−0.721	−0.849	−0.922*	−1.052**
	(0.529)	(0.431)	(0.347)	(0.312)
Log median household income (1980)	−0.072	−0.095**	−0.046	−0.079*
	(0.040)	(0.028)	(0.034)	(0.030)
Pop. density (1980)	−5.169	0.599	−7.494	−4.536
	(10.309)	(8.102)	(7.052)	(5.747)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes
R^2	0.597	0.665	0.595	0.643
Adj. R^2	0.552	0.628	0.545	0.598
Num. obs.	403	403	373	373
RMSE	2.333	2.125	2.434	2.286
N Clusters	37	37	33	33

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table D-3: Effect of proximity to diversifying tract on support for voter's veto, among precincts which changed minimally from 1980 to 1993.

E Main Results by Racial Groups

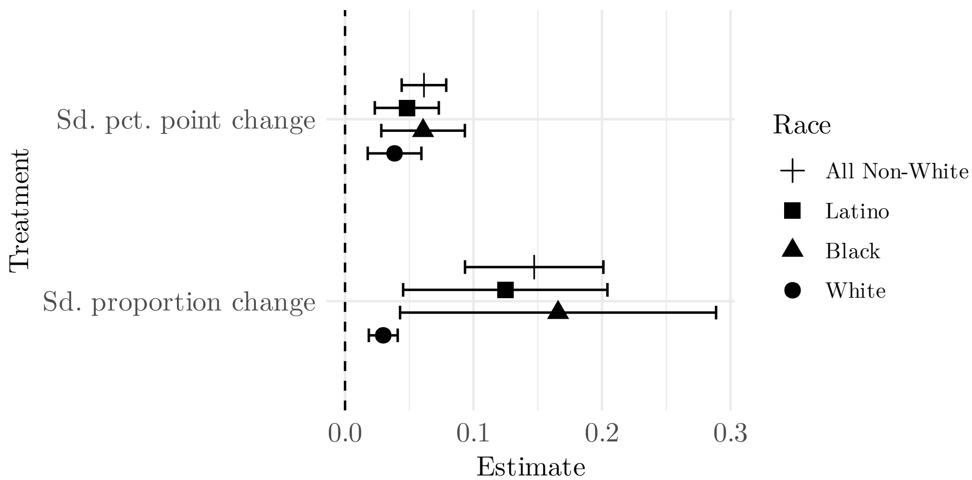


Figure E-8: Variation in the treatment effect of local racial demographic change in precincts with majority non-white, majority Black, majority Latino, and majority white residents pretreatment (1980). Point estimates shown with 95%-confidence intervals.

To assess whether this effect generalizes to majority Black and majority Latino communities, I adapt the interaction model to compare precincts with a majority of each group to precincts with a minority of the same racial group. For context, as of 1980, there were 401 majority Latino precincts and 372 majority Black precincts. However, because there were only 8 majority Asian precincts in 1980, I do not estimate an effect among majority Asian precincts.

As shown in Figure E-8, majority Black and Latino precincts show the same increase in support for the voters' veto as do majority non-white precincts as whole. In contrast, the effect in majority white precincts — while positive — is generally substantively and often statistically smaller than that of other communities.

Additionally, I estimate the effect of an increase in each minority racial group to see whether the effect is driven by a specific minority group. This approach faces empirical challenges though. An increase in one racial group may happen in conjunction with that of another group. Thus, isolating the effect of one racial group is likely to misrepresent the overall treatment of increasing a decreasing white population in these precincts.

To account for these multiple changing racial groups, I include the percentage point and proportion treatments for each minority group (non-Hispanic Black, Hispanic, Asian, and Other). I also include lagged levels of each racial group to help account for each precinct's pretreatment composition. Finally, because each of these groups are different sizes, I no longer define treatment as a standard deviation increase. Here, the percentage point change model is based on moving from 0 percent to 100 percent of that ethnic group, whereas the proportion model is defined as a 100 percent increase in group size. While these are unrealistic changes, they place all of the coefficients on the same absolute scale.

As shown in Table E-4, a percentage point increase in the population of any of the major non-white racial groups increases support for the voters veto. In contrast, the treatment effect of a proportion change seems to be driven by a proportion increase in the Latino population. Recall, the main mechanism of this study is a decrease in the local white population. While the declining white population may still be the main driver, it may also be the case that precincts losing white residents were experiencing the largest proportional increases in Latino residents, hence the connection. Still, it is difficult to identify the extent to which this is an empirical anomaly due to isolating fractions of the overall treatment (an increase in the local percent non-white).

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in Hisp. pop.	0.305^{**}	0.399^{***}		
	(0.106)	(0.067)		
Pct. point Δ in Black pop.	0.952^{***}	0.419^{**}		
	(0.147)	(0.139)		
Pct. point Δ in Asian pop.	0.402^{***}	0.243^{***}		
	(0.115)	(0.055)		
Pct. point Δ in Other pop.	2.812^{***}	0.330		
	(0.709)	(0.296)		
Prop. Δ in Hisp. pop.			-0.029^{***}	0.037^{***}
			(0.006)	(0.009)
Prop. Δ in Black pop.			0.024^{*}	0.007
			(0.011)	(0.004)
Prop. Δ in Asian pop.			0.016^{*}	0.003
			(0.007)	(0.003)
Prop. Δ in Other pop.			0.010	−0.006
			(0.007)	(0.004)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes
R^2	0.633	0.771	0.557	0.755
Adj. R^2	0.618	0.761	0.539	0.745
Num. obs.	3718	3651	3708	3642
RMSE	3.091	2.453	3.397	2.535
N Clusters	144	142	142	140

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table E-4: Effect of change in non-white population (1980-199) on support for voters' veto, disaggregated by race of change. Treatment data are winsorized.

Finally, to test whether the main effect is driven by either a concern over losing white residents or rather a decrease in one's own racial group which may stimulate group threat among minorities, I test whether majority Black and Latino precincts show a similar increase in support for the voters' veto when their own racial group decreases in size. Isolating these majority Black and majority Latino precincts, I find that a decrease in each group's own share of the population does not elicit an exclusionary reaction (Tables E-5 and E-6).

Only a decrease in the local white population spurs support for local control over affordable housing. These subgroup analyses suggest that both white and non-white voters may be feeling concerns over the economic, social, and political trajectory of their neighborhoods as the local white population shrinks, rather than concerns over the size and influence of their own racial group.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-Black pop.	-0.030^{*}	-0.003		
	(0.013)	(0.006)		
Prop. Δ in non-Black pop.			-0.029^{*}	-0.008
			(0.014)	(0.012)
Pct. unemployed (1980)		-0.272^{***}		-0.258^{***}
		(0.047)		(0.049)
Pct. homeowner (1980)		0.032		0.036
		(0.029)		(0.031)
Pct. vacant (1980)		-0.171		-0.131
		(0.192)		(0.166)
Log median household income (1980)		0.063^{***}		0.056^{***}
		(0.013)		(0.012)
Pop. density (1980)		-2.489		-2.126
		(2.799)		(2.949)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.489	0.585	0.497	0.587
Adj. R^2	0.462	0.555	0.470	0.557
Num. obs.	372	369	372	369
RMSE	2.223	2.025	2.206	2.020
N Clusters	19	19	19	19

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table E-5: Effect of change in a precinct's non-Black population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-Hisp. pop.	−0.011	−0.013		
	(0.009)	(0.007)		
Prop. Δ in non-Hisp pop.			−0.013	-0.014^{*}
			(0.007)	(0.006)
Pct. Hispanic (1980)		−0.038		−0.029
		(0.049)		(0.059)
Pct. unemployed (1980)		-0.838^{**}		-0.830^{*}
		(0.303)		(0.313)
Pct. homeowner (1980)		0.067		0.068
		(0.064)		(0.063)
Pct. vacant (1980)		0.401		0.393
		(0.331)		(0.337)
Log median household income (1980)		0.156^{***}		0.156^{***}
		(0.034)		(0.034)
Pop. density (1980)		-3.314^{***}		-3.194^{***}
		(0.823)		(0.882)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.359	0.600	0.362	0.600
Adj. R^2	0.280	0.541	0.283	0.541
Num. obs.	401	393	401	393
RMSE	4.347	3.492	4.337	3.493
N Clusters	44	44	44	44

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table E-6: Effect of change in a precinct's non-Hispanic population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized.

F Ecological Inference

To account for the challenge of aggregate precinct-level data, I use King (1997)'s Ecological Inference (EI) to estimate individual-level population behavior. The most severe challenge to the assumptions necessary for EI occurs from violations of the "no aggregation bias assumption" (Cho 1998). In the context of this study, the no aggregation bias assumption implies that white (non-white) voters in predominantly white (non-white) precincts would support the voters' veto at the same or similar rate as a white (non-white) voter in a predominantly non-white (white) precinct.

For non-white voters, this may be a reasonable assumption, given non-white voters strong support for the Democratic party at the time and the lack of heterogeneity in the majority non-white precincts in Figure B-3. However, this assumption is more questionable among white voters, both given the wide variation in support for the voters' veto in predominantly white precincts and in the ability of greater ability white voters to sort into their preferred neighborhoods at the time. Due to trends around residential segregation and sorting, it is likely that the white voters living in more racially diverse precincts were more liberal than white voters in more homogeneous precincts.

It this bias exists, then EI would underestimate the liberalness of white voters in non-white precincts and overestimate the liberalness of white voters in predominantly white precincts. Additionally, I apply weights to each precinct inversely proportional to the size of the ecological inference standard error, as suggested by Adolph et al. (2003) and King (1997). This process down-weights whites in racially diverse precincts and up-weights those in homogeneously white precincts. In short, the EI method up-weights precincts for which we may incorrectly categorize whites as more liberal than they truly are. This tendency of EI would bias my results towards finding that white voters behave in a liberal manner, such as having no reaction in response to local racial demographic change.

Instead, the EI results support the evidence presented from the aggregate data: both white and non-white voters respond to local racial demographic change by supporting the conservative, voters' veto policy. To arrive at this conclusion, I combine the precinct-level data with 1992 voter registration data from Enos, Kaufman and Sands (2019). The voter data includes estimates of voter registration composition by race and allows me to model voter behavior in over 2,000 precincts, roughly 66% of the precincts my precinct-level analysis.

Using EI, I generate vote shares for both white and non-white voters within each precinct. I then use those vote shares as the dependent variable in the otherwise same models as reported in Table 1. The first check is whether this subset of precincts behaves similarly to the full sample of 3,718 precincts. It does. Table F-7 shows the results from this subset are almost identical to the main results reported in Table 1.

Next, what is the effect of local racial demographic change on the voting behavior of both white and non-white voters? Table F-8 shows that for the two definitions of local demographic change both white and non-white voters show greater support for the voters' veto when experiencing a local demographic shift. This is additional evidence that the effects of racial demographic change are not solely driven by white voters in response to racial threat.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.065***	0.049***		
	(0.006)	(0.005)		
Prop. Δ in non-white pop.			0.074***	0.039***
			(0.007)	(0.009)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.563	0.760	0.591	0.725
Adj. R^2	0.545	0.750	0.574	0.712
Num. obs.	2172	2149	2172	2149
RMSE	3.550	2.636	3.437	2.825
N Clusters	87	87	87	87

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table F-7: Effect of change in a precinct's non-white population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized. Using subset of data merged with voter registration data.

	Support for voters' veto, 1993			
	White	Non-white	White	Non-white
Pct. point Δ in non-white pop.	0.066***	0.043***		
	(0.008)	(0.003)		
Prop. Δ in non-white pop.			0.047***	0.018**
			(0.014)	(0.006)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes
R^2	0.732	0.627	0.698	0.580
Adj. R^2	0.719	0.610	0.685	0.561
Num. obs.	2149	2149	2149	2149
RMSE	0.375	0.312	0.398	0.331
N Clusters	87	87	87	87

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table F-8: Effect of change in a precinct's non-white population (1980-199) on support for voters' veto, 1993. Treatment data are winsorized. Controls included in all models, but excluded from table due to space constraints. Using subset of data merged with voter registration data.

G Individual-Level Survey Data

One concern from the observational analysis is whether the test of heterogeneous effects by precinct-level racial majority is reflecting the behavior of members of that racial demographic group. Given exclusionary behavior in housing is commonly associated with white voters, how can we be sure that the treatment effect in majority Black precincts is not just the reaction of white voters living there? Pushing beyond Ecological Inference in Section F, is there direct evidence that non-white voters also exhibit exclusionary behavior in response to local demographic change?

I have not been able to find raw survey data that matches my period of interest (1980-1993). However, contemporary survey can help capture whether assess whether non-white residents exhibit similar behavior. I gathered as much survey data as I could find capturing the outcome of interest: the loss of local control over housing. I found three recent surveys that ask questions of this type:

1. Public Policy Institute of California (PPIC), fielded May 19 to 28, 2019, n = 1,713 adults in California (Baldassare et al. 2019): "Next, do you favor or oppose these state government proposals to provide more affordable housing in your part of California?. . . How about requiring local governments to approve a certain amount of new housing development before they can receive state funding for their local transportation projects?. . . Do you favor or oppose this proposal?"
    - Favor
    - Oppose
2. Public Policy Institute of California (PPIC), fielded June 7 to 29, 2023, n = 1,724 adults in California (Baldassare et al. 2023): "Which of the following two statements comes closer to your views — even if neither is exactly right?"
    - The state government should require local governments to build their "fair share" of new housing that is affordable for the workforce in the region
    - Local governments should decide how much and what kinds of new housing to build in their communities
3. Pew Charitable Trusts, fielded September 8 to 17, 2023, n = 5,051 adults nationwide (Horowitz and Kansal 2023): "For each, please indicate how much you support or oppose that idea. . . Require local governments to use a quick and clear process for making decisions about building permits."
    - Strongly support
    - Somewhat support
    - Somewhat oppose
    - Strongly oppose

I converted the Pew question to a binary outcome of support to match the format of the two PPIC questions. I define support for the voters' veto as support for local control over housing, compared to the state level or limitations on the local government. I then use the respondents' ZIP codes to merge the survey data with Census demographic change data. To mimic the same length of time as the observational data (14 years), I calculate the ZIP code-level change in percent non-white from 2007 to 2020. Together, this dataset contains 8,296 respondents with ZIP code-level demographic change data.

I regress support for the voters' veto on the change in percent and proportion non-white from 2007 to 2020. To assess differential effects among white and non-white respondents, I interact the treatment with a binary for respondent racial identity (1 = white, 0 = non-white). I also control for ZIP code percent white pretreatment (2007) as well as the demographic controls that may be associated with the treatment and outcome and are common among the three surveys, including income, homeownership, and education. I do not control for political party or ideology, as those were not collected for the Pew survey. I include a survey-level fixed effect to account for variation in the survey timing and question wording and use robust standard errors.

Table G-9 show the results for the percentage point change in Columns 1 and 2 and proportion change in Columns 3 and 4. A standard deviation increase in the percentage point non-white ZIP code population is associated with 2.4 to 3.1 points ($d = .06$ to $d = .08$) greater support for the voters' veto. The null interaction term indicates that this relationship is shared among white and non-white respondents. An standard deviation increase in the proportion non-white is associate with a 2.3 to 4.7 points ($d = .06$ to $d = .11$) greater support for the voters' veto. While this proportion change effect is concentrated among non-white voters, the same is true for the observational data when change is defined by a proportion change (Table 2).

These individual survey results match the pattern of the observational data and support the evidence that the voters' veto is shared among both white and non-white voters. Of course, it is possible that many of these respondents may have moved into the area as part of the demographic change from 2007 to 2020. To address this concern, I replicate the analysis using only homeowners, who are more likely to be residentially stable and have experienced a longer duration of this demographic change. Likewise, homeowners mortgages are more likely to keep them tied to their location, even if they dislike the changes they are experiencing, limiting the risk of selection out of the context.

Table G-10 shows that the results using only homeowners replicate the pattern from the observational analysis. Exclusionary responses to demographic change are as great if not greater among non-white respondents as they are among white respondents.

	Support for voters' veto			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.031^{***}	0.024^{***}		
	(0.007)	(0.007)		
Prop. Δ in non-white pop.			0.047^{***}	0.023
			(0.010)	(0.012)
Pct. point Δ x white resp.	-0.009	-0.006		
	(0.009)	(0.009)		
Prop. Δ x white resp.			-0.056^{***}	-0.041^{***}
			(0.011)	(0.011)
White, non-Hispanic	0.033^{***}	0.012	0.023^{*}	0.003
	(0.010)	(0.011)	(0.010)	(0.011)
Pct. white, non-Hisp. (2007)		0.017		0.051
		(0.021)		(0.028)
Income (+$75k)		0.027^{**}		0.028^{**}
		(0.010)		(0.010)
Homeowner		0.057^{***}		0.061^{***}
		(0.011)		(0.011)
College graduate		-0.029^{**}		-0.031^{***}
		(0.010)		(0.010)
Survey FE	Yes	Yes	Yes	Yes
R^2	0.105	0.111	0.103	0.110
Adj. R^2	0.104	0.110	0.103	0.109
Num. obs.	8004	7831	8004	7831
RMSE	0.391	0.388	0.392	0.388

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table G-9: Effect of change in percent non-white population (2007-20) on support for voters' veto, using survey data.

	Support for voters' veto			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.035^{***}	0.032^{***}		
	(0.009)	(0.010)		
Prop. Δ in non-white pop.			0.049^{***}	0.027
			(0.013)	(0.016)
Pct. point Δ x white resp.	-0.019	-0.016		
	(0.011)	(0.012)		
Prop. Δ x white resp.			-0.058^{***}	-0.044^{**}
			(0.014)	(0.015)
White, non-Hispanic	0.024	0.009	0.006	-0.005
	(0.012)	(0.014)	(0.013)	(0.014)
Pct. white, non-Hisp. (2007)		0.043		0.075^{*}
		(0.026)		(0.034)
Income (+$75k)		0.029^{*}		0.031^{*}
		(0.012)		(0.012)
College graduate		-0.024^{*}		-0.025^{*}
		(0.011)		(0.011)
Survey FE	Yes	Yes	Yes	Yes
R^2	0.158	0.155	0.157	0.154
Adj. R^2	0.157	0.154	0.156	0.153
Num. obs.	5330	5260	5330	5260
RMSE	0.384	0.381	0.384	0.381

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table G-10: Effect of change in percent non-white population (2007-20) on support for voters' veto, using survey data among homeowners.

H Spatial Durbin Error Model

Spatial autocorrelation can be a problem when using geographic data due to the data's violation of SUTVA. Neighboring observations can influence each other, violating SUTVA's independence assumption. Likewise, omitted variables may be spatially correlated. Fortunately, these relationships can be modeled.

Using a spatial model requires two steps: the definition of spatial relationships and the choice of a model to account for spatial dependence. The most common way to model spatial relationships is by using queen contiguity weights. These weights account for spatial interactions between units sharing either a border or a vertex. Contiguity weights are especially helpful for the precinct data used in this analysis because the relationships are not sensitive to absolute distance. For example, some parts of Los Angeles County are very dense, where spillovers may only extend 500 meters. In contrast, interactions between precincts in rural Los Angeles County may occur at the level of several kilometers. By defining spatial interactions at the level of shared precinct borders (i.e., a queen continguity weights matrix), this design accounts for variation in the scale of human interaction across Los Angeles County.

To select a model, I considered how spatial dependence may be affecting the analysis. First, the treatment of demographic change is likely to be affecting neighboring precincts. In fact, the second analysis of this study is based on exposure to nearby diversifying precincts. Thus, the model would need to account for this spillover effect. Second, there are likely omitted variables which may be spatially correlated. A spatial Durbin error model (SDEM) accounts for the spatial spillover effects of covariates as well as autocorrelation in the error term (LeSage 2014).

I reproduce the results from the Table 1 using the SDEM. Due to the computational intensity of incorporating the spatial autocorrelation, adjustments were made to the model specifications. Models 1 and 3 include Census place-level fixed effects, but no covariates. Models 2 and 4 include covariates, but no fixed effects. The covariates in Models 2 and 4 are the same as those in the non-spatial results: the 1980 precinct-level percent non-white residents, homeownership rate, vacancy rate, population density, log median household income, and unemployment rate. The results are substantively similar across all four models. All models use the queen continuity weights matrix and are weighted by precinct population.

Having accounted for spatial spillovers and spatially correlated error terms, a standard deviation increase in the local non-white population increases support for direct democracy over housing by 2.0 to 2.2 percentage points. When defined as a proportion increase, the effect is a 0.8 to 2.8 point increase in support for the voters' veto over affordable housing. These effects substantively match the results from the non-spatial OLS model in Table 1.

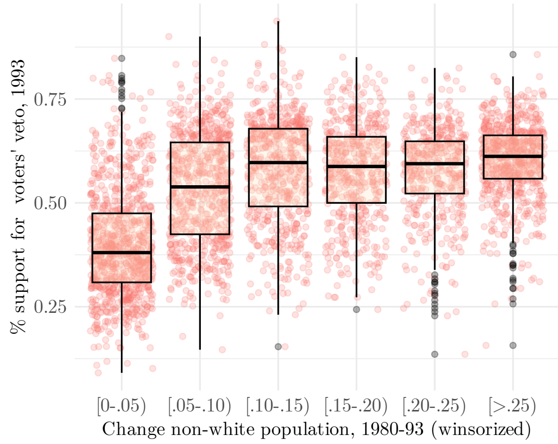
	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.022^{***}	0.020^{***}		
	(0.003)	(0.003)		
Prop. Δ in non-white pop.			0.028^{***}	0.008^{*}
			(0.003)	(0.003)
Pct. white, non-Hispanic (1980)		0.099^{***}		0.110^{***}
		(0.014)		(0.015)
Pct. unemployed (1980)		-0.581^{***}		-0.644^{***}
		(0.074)		(0.076)
Pct. homeowner (1980)		0.095^{***}		0.084^{***}
		(0.012)		(0.012)
Pct. vacant (1980)		0.072		0.079
		(0.068)		(0.069)
Log med. h.h. income (1980)		0.013		0.009
		(0.008)		(0.009)
Pop. density (1980)		-3.178^{***}		-3.197^{***}
		(0.922)		(0.938)
Place Fixed Effects	Yes	No	Yes	No
Controls	No	Yes	No	Yes
Num. obs.	3112	3112	3112	3112
Parameters	235	17	235	17
Log Likelihood	3592.387	3653.932	3637.647	3602.665
AIC (Linear model)	−5404.983	−6233.041	−5852.253	−5851.342
AIC (Spatial model)	−6714.774	−7273.864	−6805.294	−7171.331
LR test: statistic	1311.790	1042.823	955.041	1321.988
LR test: p-value	0.000	0.000	0.000	0.000

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

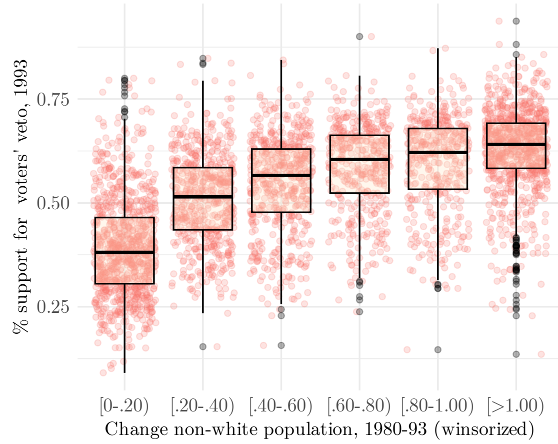
Table H-11: Effect of change in non-white population (1980-199) on support for voters' veto, spatial Durbin error model. Treatment data are winsorized.

I Visualization of Bivariate Relationships

Figure I-9 shows the bivariate relationship between my two treatment variables and precinct support for the voters' veto (Article 34 repeal) using binned box and whiskers plots. Precincts that experienced greater increases in non-white population from 1980 to 1993 showed greater support for the voters' veto over affordable housing in the 1993 November election.

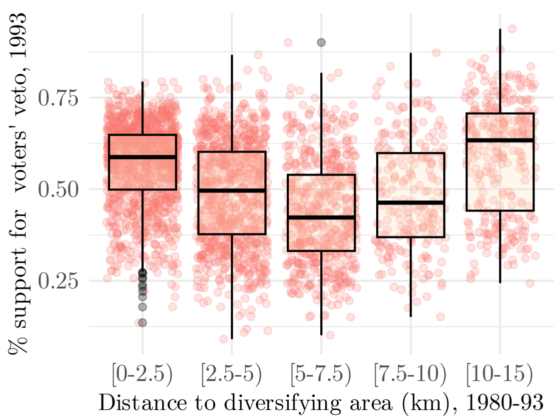


(a) Percentage point change

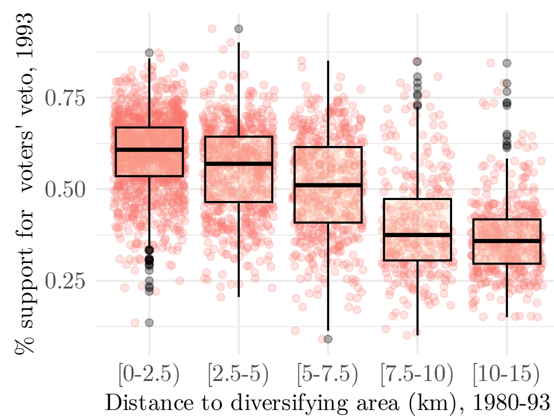


(b) Proportion change

Figure I-9: Relationship between precinct-level change in non-white population (winsorized) and support for Article 34 repeal.



(a) Percentage point change



(b) Proportion change

Figure I-10: Relationship between distance from a diversifying Census tract and support for Article 34 repeal in 1993. 10 to 15 kilometers binned together due to data sparseness.

Figure I-10 shows the bivariate relationship between a precinct's distance from a stimulus tract and support for the voters' veto using binned box and whiskers plots. Precincts farther from these diversifying tracts — defined using either percentage point change or proportion change — show less support for direct democracy over housing.

J Alternative Explanations

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in pop. density, 1980-93	-0.029^{***}	0.010^{*}		
	(0.003)	(0.005)		
Prop. Δ in pop. density, 1980-93			-0.006	0.015
			(0.007)	(0.009)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.535	0.713	0.497	0.713
Adj. R^2	0.516	0.701	0.477	0.701
Num. obs.	3718	3651	3718	3651
RMSE	3.481	2.744	3.619	2.742
N Clusters	144	142	144	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table J-12: Effect of change in a precinct's population density on support for voters' veto, 1993. Treatment data are winsorized. Controls included in all models, but excluded from table due to space constraints.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in new arrivals, 1980-93	-0.020^{*}	0.001		
	(0.008)	(0.002)		
Prop. Δ in new arrivals, 1980-93			-0.024^{**}	0.002
			(0.008)	(0.002)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.506	0.710	0.511	0.710
Adj. R^2	0.486	0.698	0.491	0.698
Num. obs.	3684	3651	3684	3651
RMSE	3.591	2.758	3.572	2.758
N Clusters	143	142	143	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table J-13: Effect of change the rate of residential churn within the past 10 years on support for voters' veto, 1993. Treatment data are winsorized. Controls included in all models, but excluded from table due to space constraints.

Regarding Table J-15, the null effect may be due to two sources of measurement error. First, during the period of this study, there was considerable debate among demographers

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in med. hh. income, 1980-93	0.054^{***}	-0.030^{*}		
	(0.003)	(0.015)		
Prop. Δ in med. hh. income, 1980-93			-0.008	-0.016^{*}
			(0.004)	(0.007)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.569	0.719	0.496	0.717
Adj. R^2	0.551	0.707	0.475	0.705
Num. obs.	3664	3649	3664	3649
RMSE	3.357	2.715	3.631	2.723
N Clusters	142	142	142	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table J-14: Effect of change in a precinct's median household income on support for voters' veto, 1993. Treatment data are winsorized. Controls included in all models, but excluded from table due to space constraints.

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in foreign born, 1980-93	-0.005	0.015^{***}		
	(0.006)	(0.003)		
Prop. Δ in foreign born, 1980-93			-0.000	0.002
			(0.003)	(0.002)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.497	0.717	0.495	0.710
Adj. R^2	0.477	0.705	0.475	0.698
Num. obs.	3718	3651	3685	3651
RMSE	3.619	2.725	3.630	2.757
N Clusters	144	142	144	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table J-15: Effect of change in a precinct's percent foreign born on support for voters' veto, 1993. Treatment data are winsorized. Controls included in all models, but excluded from table due to space constraints.

over how to accurately measure the percent foreign born, let alone the share of immigrants who were non-white (Van Hook et al. 2006; Woodrow-Lafield 1995). Second, percent foreign born was measured via the long-form Census which only sampled 1 in 6 households. While this sampling approach may be sufficient for states and places, it faces limits when aggregated to the precincts with a median size of roughly 1,200 residents. In contrast, these limitations do not exist for the percent non-white population data, which comes from the short-form, full-population Census.

K Stability Across Model Specifications

	Support for voters' veto, 1993		
	Model 1	Model 2	Model 3
Pct. point Δ in non-white pop.	0.070***	0.046***	0.044***
	(0.009)	(0.006)	(0.007)
Pct. white, non-Hispanic (1980)		0.137**	0.151**
		(0.043)	(0.052)
Pct. unemployed (1980)		−0.633***	−0.851***
		(0.154)	(0.183)
Pct. homeowner (1980)		0.198***	0.237***
		(0.033)	(0.032)
Pct. vacant (1980)		0.016	−0.299
		(0.172)	(0.211)
Log median household income (1980)		−0.042**	−0.065***
		(0.015)	(0.012)
Pop. density (1980)		−7.922***	−7.169**
		(1.595)	(2.215)
Δ pct. unemployed (1980)			−0.356*
			(0.143)
Δ pct. homeowner (1980)			0.158***
			(0.035)
Δ pct. vacant (1980)			−0.505**
			(0.171)
Δ log median household income (1980)			−0.029**
			(0.010)
Δ pop. density (1980)			−0.928
			(3.417)
Place Fixed Effects	No	No	No
Num. obs.	3718	3651	3647
N Clusters	144	142	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table K-16: Effect of change in percent non-white population (1980-1993) on support for voters' veto, removing fixed effects and varying controls. Treatment data are winsorized.

	Support for voters' veto, 1993		
	Model 1	Model 2	Model 3
Prop. Δ in non-white pop.	0.089^{***}	0.045^{***}	0.038^{***}
	(0.004)	(0.006)	(0.006)
Pct. white, non-Hispanic (1980)		0.071	0.110^{*}
		(0.039)	(0.043)
Pct. unemployed (1980)		-0.944^{***}	-1.154^{***}
		(0.163)	(0.208)
Pct. homeowner (1980)		0.195^{***}	0.232^{***}
		(0.034)	(0.032)
Pct. vacant (1980)		0.023	-0.530^{*}
		(0.173)	(0.236)
Log median household income (1980)		-0.054^{***}	-0.072^{***}
		(0.014)	(0.012)
Pop. density (1980)		-7.666^{***}	-9.572^{***}
		(2.018)	(2.143)
Δ pct. unemployed (1980)			-0.414^{**}
			(0.150)
Δ pct. homeowner (1980)			0.148^{***}
			(0.038)
Δ pct. vacant (1980)			-0.816^{***}
			(0.188)
Δ log median household income (1980)			-0.047^{***}
			(0.009)
Δ pop. density (1980)			7.132^{*}
			(3.128)
Place Fixed Effects	No	No	No
Num. obs.	3718	3651	3647
N Clusters	144	142	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table K-17: Effect of change in percent non-white population (1980-1993) on support for voters' veto, removing fixed effects and varying controls. Treatment data are winsorized.

L Internal Change, Non-Winsorized Data

	Support for voters' veto, 1993			
	Model 1	Model 2	Model 3	Model 4
Pct. point Δ in non-white pop.	0.051***	0.042***		
	(0.013)	(0.010)		
Prop. Δ in non-white pop.			0.070***	0.030**
			(0.011)	(0.011)
Pct. white, non-Hispanic (1980)		0.071*		0.073
		(0.034)		(0.041)
Pct. unemployed (1980)		−0.546***		−0.673**
		(0.156)		(0.228)
Pct. homeowner (1980)		0.137***		0.124**
		(0.039)		(0.039)
Pct. vacant (1980)		−0.307**		−0.277*
		(0.112)		(0.114)
Log median household income (1980)		0.007		−0.007
		(0.015)		(0.016)
Pop. density (1980)		−6.017***		−5.470***
		(0.533)		(0.780)
City FE	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
R^2	0.568	0.747	0.601	0.720
Adj. R^2	0.550	0.736	0.584	0.708
Num. obs.	3718	3651	3718	3651
RMSE	3.355	2.575	3.225	2.711
N Clusters	144	142	144	142

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table L-18: Effect of change in percent nonwhite population (1980-199) on support for voters' veto. Treatment data are not winsorized.

M Proximity to Change, Alternative Percentile Thresholds

	Proportion Change		
	Model 1	Model 2	Model 3
Proximity to change (95th)			0.020
			(0.014)
Proximity to change (90th)		0.030^{***}	
		(0.008)	
Proximity to change (85th)	0.030^{***}		
	(0.008)		
Place Fixed Effects	Yes	Yes	Yes
Controls	Yes	Yes	Yes
R^2	0.594	0.597	0.576
Adj. R^2	0.549	0.552	0.530
Num. obs.	403	403	403
RMSE	2.340	2.333	2.391
N Clusters	41	37	27

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table M-19: Effect of proximity to diversifying tract on support for voter's veto. Controls included in all models, but excluded from table due to space constraints.

	Proportion Change		
	Model 1	Model 2	Model 3
Proximity to proportion change (95th)			0.058***
			(0.008)
Proximity to proportion change (90th)		0.052***	
		(0.008)	
Proximity to proportion change (85th)	0.022		
	(0.012)		
Place Fixed Effects	Yes	Yes	Yes
Controls	Yes	Yes	Yes
R^2	0.584	0.665	0.682
Adj. R^2	0.538	0.628	0.647
Num. obs.	403	403	403
RMSE	2.370	2.125	2.072
N Clusters	41	37	27

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table M-20: Effect of proximity to diversifying tract on support for voter's veto. Controls included in all models, but excluded from table due to space constraints.

	Proportion Change		
	Model 1	Model 2	Model 3
Proximity to change (95th)			0.003
			(0.009)
Proximity to change (90th)		0.018**	
		(0.006)	
Proximity to change (85th)	0.017**		
	(0.006)		
Place Fixed Effects	Yes	Yes	Yes
Controls	Yes	Yes	Yes
R^2	0.591	0.595	0.581
Adj. R^2	0.541	0.545	0.529
Num. obs.	373	373	373
RMSE	2.445	2.434	2.476
N Clusters	38	33	22

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table M-21: Effect of proximity to diversifying tract on support for voter's veto. Controls included in all models, but excluded from table due to space constraints.

	Proportion Change		
	Model 1	Model 2	Model 3
Proximity to proportion change (95th)			0.044***
			(0.007)
Proximity to proportion change (90th)		0.040***	
		(0.007)	
Proximity to proportion change (85th)	0.015		
	(0.011)		
Place Fixed Effects	Yes	Yes	Yes
Controls	Yes	Yes	Yes
R^2	0.591	0.643	0.659
Adj. R^2	0.540	0.598	0.617
Num. obs.	373	373	373
RMSE	2.446	2.286	2.233
N Clusters	38	33	22

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table M-22: Effect of proximity to diversifying tract on support for voter's veto. Controls included in all models, but excluded from table due to space constraints.

N Proximity to Change, by 1980 Racial Composition

This analysis is conducted on precincts in the top tercile percent white (> 85%) and the bottom tercile percent white (< 52%). These values differ slightly from the "within precinct" analysis as the data are subset to precincts which changed only minimally between 1980 and 1993. Proximity to a diversifying tract is interacted with a binary variable for whether the precinct was in the top percent white tercile in 1980.

Results here are noisier due to the very limited sample sizes. However, Precincts that were in the bottom tercile percent white still show a generally statistically significant reaction to the proximity of a diversifying tract. The lower order term in Model 1 is significant at the $\alpha = 0.10$.

	Percentage Point Δ		Proportion Δ	
	Model 1	Model 2	Model 3	Model 4
Prox. to pct. point Δ in non-white	0.028		0.010	
	(0.017)		(0.009)	
Prox. to prop. Δ in non-white		0.056^{***}		0.043^{***}
		(0.011)		(0.008)
Prox. to pct. point x white prec., 1980	0.005		−0.002	
	(0.023)		(0.019)	
Prox. prop. x white prec., 1980		−0.008		-0.049^{*}
		(0.017)		(0.019)
Pct. unemployed (1980)	−0.472	-0.977^{***}	-0.909^{***}	-1.283^{***}
	(0.255)	(0.210)	(0.198)	(0.191)
Pct. homeowner (1980)	0.026	0.034	−0.057	−0.058
	(0.052)	(0.052)	(0.047)	(0.039)
Pct. vacant (1980)	−1.068	-0.964^{*}	-0.873^{*}	-0.725^{*}
	(0.605)	(0.451)	(0.349)	(0.309)
Log median household income (1980)	−0.047	-0.074^{*}	0.023	−0.018
	(0.043)	(0.033)	(0.029)	(0.024)
Pop. density (1980)	−8.927	−3.601	−13.427	−10.552
	(11.408)	(8.387)	(7.109)	(5.881)
Place Fixed Effects	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes
R^2	0.599	0.664	0.596	0.654
Adj. R^2	0.549	0.622	0.539	0.605
Num. obs.	374	374	340	340
RMSE	2.313	2.118	2.447	2.266
N Clusters	36	36	32	32

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table N-23: Effect of proximity to diversifying tract on support for voter's veto, among precincts which changed minimally from 1980 to 1993.

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