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SINGLE-FAMILY ZONING AND RACE
EVIDENCE FROM THE TWIN CITIES

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Abstract

A new wave of activism and research has renewed critiques of single-family zoning as a means of racial exclusion. To test this argument, we assemble digital zoning data covering the Minneapolis–St. Paul metro area and quantify the relationship between different types of residential zoning and racial and ethnic shares of neighborhood populations. We find that a neighborhood zoned for multifamily housing has a non-White population share that is 21 percentage points higher than that of an equally situated single-family neighborhood. By contrast, we find relatively modest differences in racial shares across neighborhoods with differing minimum lot sizes. We argue that these patterns are explained by racial homeownership differences, which are especially severe in the Twin Cities. Policymakers pursuing racial integration should take into account racial differences in tenure and ensure that housing types suited for both owner and rental occupancy are allowed in all neighborhoods.

JEL codes: R52, J15, R21

Keywords: zoning, middle housing, segregation, Black homeownership, homeownership gap, housing tenure, rental housing, Minneapolis, Twin Cities, Minnesota

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Single-Family Zoning and Race: Evidence from the Twin Cities

Salim Furth and MaryJo Webster

1. Introduction

Recent scholarship has renewed criticism of the use of restrictive zoning as a successor to institutions of racial residential exclusion such as restrictive covenants.[1] Advocates for zoning reform, notably including the group Desegregate Connecticut, have highlighted zoning's role in racial exclusion.[2] However, few researchers have tried to quantify the relationship between zoning and race at a neighborhood level[3] or to distinguish among different types of land use restriction.

This paper is among the first to fill that gap, verifying and quantifying the relationship between zoning typology and minimum lot size and the racial residential patterns in Minnesota's Twin Cities metro area. We find that if a neighborhood's residential land is zoned for multifamily housing, it has, on average, a non-White population share that is 21 percentage points higher than that of an equally situated neighborhood zoned for single-family housing. Zoning for middle housing,[4] such as duplexes, yields a relationship two-thirds as strong. And zoning for Planned Unit Developments (PUDs), which often include a mix of housing typologies, is also associated with a higher non-White population share. We also find some evidence that densely zoned single-family areas have higher non-White population shares, but that result is less robust.

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[1] See, for example, Rothstein (2017), Shertzer et al. (2022), and Whittemore (2017). Earlier critics include Babcock and Bosselman (1963), Davidoff and Gold (1970), Mandelkar (1976), and Pendall (2000), among many others.

[2] See, for example, Demsas (2021).

[3] Our neighborhoods are census block groups, which typically contain 600 to 3,000 residents.

[4] Wegmann (2020) describes middle housing as "low-rise, middle-density housing" (p. 114).

We then tie our novel findings to well-established research on race and homeownership to argue that zoning's relationship to race is mediated through typology and tenure. In short, zoning guides what housing typologies are built, each typology has a strong tendency toward either owner or renter occupancy, and homeownership rates differ sharply by race. We offer evidence against competing explanations for the relationship between zoning and race, such as cultural differences in typology preference or the unique ethnic makeup of Minnesota's non-White population.

Section 2 provides a brief review of the literature most directly related to our research question. Section 3 provides background context on the racial demographics and land use regulatory institutions of our region of study. Section 4 introduces our data sources. Section 5 describes our approach to the key empirical questions of this study. Section 6 walks through the results, describing the distinctions between key regressions and summarizing their results. Section 7 discusses the channels through which the relationship between zoning and race is likely mediated and addresses some secondary questions. Section 8 concludes.

We have interspersed in our text a number of figures and a selection of summary tables that are useful reading aides. The remaining tables, including those reporting detailed analysis, appear in an appendix following the text.

2. Literature Review

Our study is situated in the small empirical corner of the intersection of the vast scholarly literatures on residential segregation and zoning. Since the earliest days of zoning, critics have pointed out its exclusionary intent. Judge David C. Westenhaver saw the single-family zoning of Euclid, Ohio, as a means to class segregation:[5]

[5] Westenhaver's decision was, of course, reversed by the Supreme Court.

> In the last analysis, the result to be accomplished is to classify the population and segregate them according to their income or situation in life. (*Ambler v. Euclid*, 1924).

Following the Civil Rights movement, the exclusionary aspect of zoning became clearly tied to race as racially explicit legal barriers were dismantled. By 1970, Davidoff and Gold could point to zoning as a primary legal barrier to integration:

> Exclusionary zoning is largely responsible for the fact that segregation by race and economic class has, over the past few decades, become accepted social policy in large metropolitan areas around the nation. (p. 60)

To the voluminous qualitative literature on the race-conscious motives and racial effects of zoning a much smaller empirical literature has recently been appended. Until recently, empirical work documenting the relationship between race and zoning has relied on surveys of city planners.

Pendall (2000) hypothesizes that adequate "new housing supply, multifamily housing supply, rental housing supply, and affordable rental housing" are important to the inclusion of non-White residents (p. 127) and that several land use controls, including "low-density" zoning,[6] degrade those supply conditions. He tests his hypotheses using data from his own 1994 survey of 1,510 jurisdictions and documents a "chain of exclusion," whereby low-density zoning lowers housing production, especially of multifamily housing, which in turn lowers the rental share and rental affordability, thereby inhibiting the growth of the Hispanic (and perhaps Black) population. Rothwell and Massey (2009) use Pendall's regulatory data to update his findings through the 2000 Census.

Using the 2008 Wharton Residential Land Use Regulatory Index survey, Lens and Monkkonen (2016) investigate the relationship between density restrictions and income

[6] Pendall interprets low density as anything less than eight units per acre.

segregation across 95 metro areas. They find that density restrictions are associated with a concentration of affluence, but not poverty. Trounstine (2020) finds correlations between the White population share and regulatory stringency at the city level; regrettably, her findings appear to suffer from omitted variable bias.

Menendian et al. (2020, 2022) relate city-level zoning to racial shares and segregation metrics in California's largest metro areas. They find relatively small differences between groups of cities that have even a modest share of multifamily zoning. But they find sharply higher shares of White residents in cities where at least 90% of residential land is limited to single-family homes.

Like the present paper, these cross-sectional studies cannot be interpreted as causal estimates.

The digitization of zoning maps has enabled a new generation of studies, to which the present paper belongs. Resseger (2022) uses block-level zoning data from Greater Boston and, to achieve causal identification, limits consideration to adjacent blocks. He finds that multifamily blocks have a Black and Hispanic population share 9 percentage points higher than do adjacent single-family blocks. At boundaries between single-family districts of different densities, he finds a small, significant effect for Hispanic but not Black residents (pp. 28–29, post-peer draft).

In ongoing research, Kulka (2021) uses a similar boundary identification strategy to study the impacts of density zoning on income and race in Wake County, North Carolina. Her data sources limit her to comparing single-family districts. She estimates that the White population will be 5 percentage points higher in a census tract that allows four fewer houses per acre (p. 21).

Our paper is less ambitious than these efforts, but it is grounded in equally good data. We add to the small and geographically specific body of empirical evidence on the contemporary relationship between zoning and race. We also draw on other data and scholarship to interpret the new findings in a framework similar to Pendall's (2000) "chain of exclusion." Helpfully, and unlike the work of Resseger (2022) and Kulka (2021), our methods can be readily replicated for other metro areas as the National Zoning Atlas takes shape.[7]

A large body of national research has shown that White families are more likely to attain and retain homeownership. Boehm and Schlottmann (2004, p. 123) show that White families are more likely to transition from renting to owning and are more likely to move up the housing ladder. Charles and Hurst (2002) show that Black renters are less likely than their White counterparts to transition to homeownership, primarily because they are less likely to apply for mortgages and secondarily because they face a much higher mortgage rejection rate.

Apart from income, the factors of credit data and parental wealth—both of which are difficult to observe—are among the most likely influences in racial gaps in homeownership. Using proprietary data sources, Dey and Brown (2022) document the former and Brandsaas (2021) documents the latter. Hilber and Liu (2008) use intergenerational wealth data from the Panel Study of Income Dynamics to argue that the residual Black–White homeownership gap can be fully accounted for by wealth and locational preferences.

3. Context

In this section, we offer a brief overview of the racial demographics of the Twin Cities region, its land use institutions, and scholarship on segregation and affordability in the region.

[7] The National Zoning Atlas is a data collection project created and led by Sara Bronin. At the time of writing, efforts are under way in 10 states, but data are not yet publicly available (https://www.zoningatlas.org/).

Racial Demographics

Until 1970, the Twin Cities population was less than 2% Black (U.S. Department of Commerce, 1974, Table A, p. iii). Far fewer Black migrants made their way to Minneapolis or St. Paul than to peer cities during the postwar Great Migration (U.S. Census Bureau, 2012).[8] Early Urban League leaders in the region, according to Delton (2001), agreed to discourage Black migration from the South in exchange for expanded employment opportunities. In 1980, the metro had only 12,700 Southern-born Black residents, compared to 64,100 in Milwaukee, 31,900 in Denver, and over half a million in Chicago (Gregory, n.d.).

Black Americans who lived in the Twin Cities during the early and mid-20th century faced significant legal and social barriers to residential integration—and the threat of violence. Crowds of thousands of White residents protested and hurled rocks for several days when Arthur and Edith Lee, a Black couple, moved to all-White Columbus Avenue in south Minneapolis (Walter, 2020).

Sood and Ehrman-Solberg (2022) use Hennepin County data from the Mapping Prejudice Project to show that racial covenants, which barred non-White residents, blocked Black residents from some new subdivisions, especially in the 1930s and 1940s (p. 41). Those covenants cast a long shadow on housing markets even after the U.S. Supreme Court ruled them unenforceable in 1948. Sood and Ehrman-Solberg find that the presence of even a small number of formerly covenanted homes has an outsize influence on the size of the 2010 Black population, perhaps indicating the informal pressure that kept neighborhoods segregated long after legal barriers were breached. Sood and Ehrman-Solberg find that the existence of a historical covenant is associated

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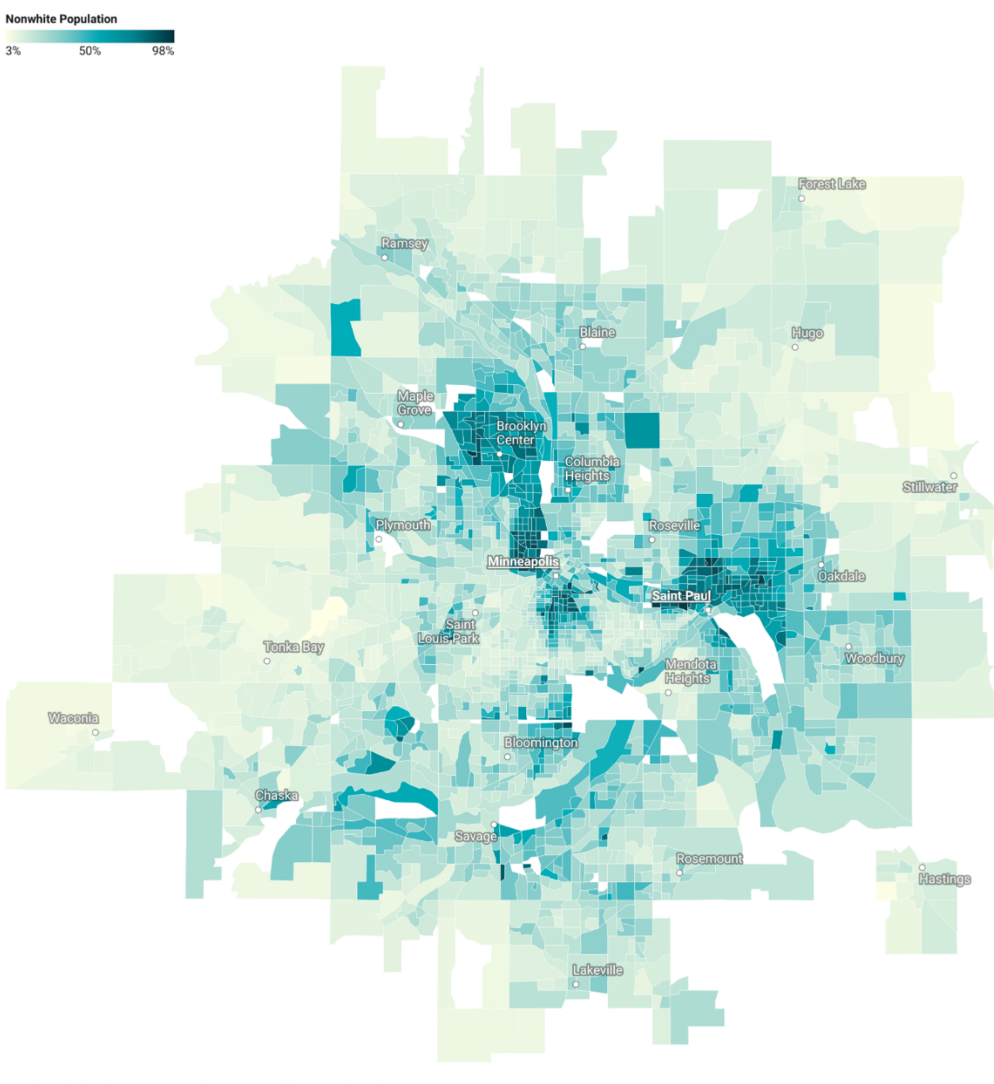
[8] The Census Bureau notes that the Black population share of the Twin Cities grew 3.4 percentage points during the smaller "First Great Migration," 1910–1940, but only 0.1 percentage points during the larger "Second Great Migration," 1940–1970.

with a 15% higher current home value. They also show an outsize effect of covenants on the modern Black population—"a 1% increase in covenanted houses in a census [block] reduces black residents by 14% and reduces black homeownership by 19%" (p. 1).

In recent decades, the non-White population has grown rapidly and has suburbanized. As of 2020, 31% of residents were non-White and 66% of non-White residents lived in the suburbs, a pattern visible in Figure 1. Orfield and Luce (2013) have documented this trend nationally.

Figure 1: Concentrated Non-White Population in the Twin Cities



The growth in non-White population includes both foreign and domestic migrants. In the 2010s over half of African American Minnesota residents had been born in another state, reflecting ongoing migration (Minnesota State Demographic Center, 2018, p. 10).

Minnesota has large immigrant communities from a handful of origins. A 2018 Minnesota State Demographic Center study of 2012–2016 American Community Survey data identified the largest cultural groups with recent immigrant histories[9] as Mexican (3.3% of the state population), Hmong (1.4%), Somali (0.9%), and Asian Indian (0.8%).

Figure 2 shows households by race of the household head and immigrant presence for the Twin Cities and several comparable metro areas. Note that the White, nonimmigrant share is especially high in the Twin Cities and that there are relatively few non-White households without an immigrant member.

Among households headed by Black residents in the Twin Cities, 47% include at least one person born abroad, compared to 15% nationwide. Statewide data indicate that Minnesotans from African immigrant communities have roughly comparable socioeconomic outcomes as African-American Minnesotans, including a comparable rate of homeownership (Minnesota State Demographic Center, 2018, p. 29).

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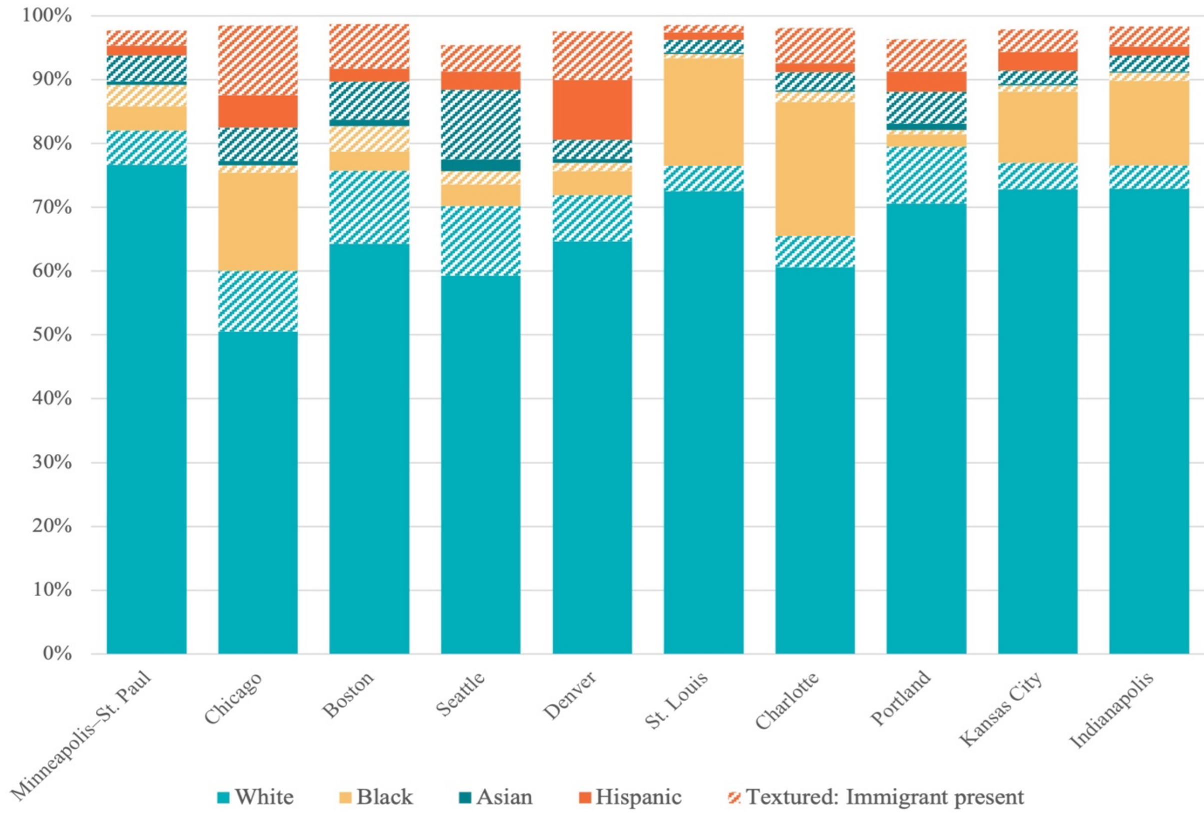
[9] We distinguish these groups from longer-established cultural groups, such as Russians, by their having at least one-third of members foreign born.

Figure 2: Households by Race and Immigrant Presence in Selected Metro Areas



Sources: American Community Survey (ACS) 2020 5-year data; Ruggles et al. (2021).

Note: Households are categorized in the figure by the racial identity of the household head and the presence of a foreign-born household member. Note that each household member may be of a different race or ethnicity.

Asian cultural and ancestry groups are split between Indians, Chinese, and Koreans, who have socioeconomic outcomes equal to or above Whites, and Southeast Asians and Filipinos, who exhibit higher rates of poverty and lower per capita earnings. However, the same groups have exceptionally high shares of multiple-earner households. Perhaps as a consequence, the homeownership rate for the Hmong—probably the least-advantaged group—is equal to that of Indians, who have the highest educational attainment and earnings (Minnesota State Demographic Center, 2018, pp. 22–29).

Institutions of Land Use Regulation

Zoning authority is exercised in Minnesota by cities and towns.[10] Counties also have the authority to regulate land use, although this power is denied to the state's two entirely urban counties, Hennepin and Ramsey (Planning, Development, Zoning, 2021).

Like other states, Minnesota adopted legislation that is based on the model planning and zoning enabling acts promoted by the U.S. Department of Commerce in the 1920s. Most cities and villages in Minnesota were authorized to adopt zoning ordinances by 1929. Counties and townships received zoning powers a decade later. These powers were reframed by the 1959 County Planning Act and the 1965 Municipal Planning Act. Townships were authorized to use the Municipal Planning Act in 1982.

The Metropolitan Land Use Planning Act (LUPA) of 1976 transformed planning and zoning for cities and townships in what is commonly called "the metro." LUPA requires regional comprehensive planning and regulation. (The metro contains almost all of seven counties, making it smaller than the Minneapolis–St. Paul–Bloomington Metropolitan Statistical Area as defined by the U.S. Office of Management and Budget.)

The Metropolitan Council of Minneapolis–St. Paul, or "Met Council," is a regional policy-making body and planning agency with board members appointed by the governor and is responsible for implementing LUPA. The Met Council also operates the region's sewer and transit systems. Its authority over sewer extension gives the Met Council practical as well as legal power to compel municipalities to enact zoning regulations that comply, at least loosely, with Met Council plans (Haigh, 2013, p. 167).

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[10] In Minnesota law, "'town' is the correct term to refer to an organized governmental unit and 'township' is the term to refer to the geographic area." Both terms are used in common speech (Minnesota House Research Department, 2019).

In the late 1970s, the then-powerful Met Council delivered notable progress in suburbanizing low- and moderate-income housing under a “fair share” system. A decade later, with less political will and fewer federal dollars to share, the Met Council essentially abandoned its fair share mandate (Goetz et al., 2003). In recent decades, the Met Council has sought consensus with municipalities, with land use issues leading to open conflict only occasionally (Mondale & Fulton, 2003).

Orfield and others have provocatively argued that the early progress toward suburban integration in Minnesota was undermined by the “Poverty Housing Industry,” which shifted from a focus on spreading affordable housing around the region to developing easier, cheaper sites within existing poor neighborhoods (Orfield et al., 2015; Orfield & Stancil, 2017).

The cities of Minneapolis and St. Paul have recently taken nationally noted measures to address housing scarcity and unaffordability. Minneapolis upzoned all of its single-family zones to middle housing (Kahlenberg, 2019); St. Paul instituted a strict rent control ordinance (Ahern & Giacoletti, 2022). However, those policies postdate our data.

4. Data

We are the first researchers to collect and exploit digitized zoning data for the Twin Cities metro. To that, we join data from conventional sources; this section describes the first at length and the latter briefly.

We requested geographic information system (GIS) shapefiles showing zoning designations by parcel from 102 municipalities in the Twin Cities.[11] The cities we chose are inside the

[11] With Michael Corey, author MaryJo Webster initially gathered and analyzed this data for a *Minneapolis Star Tribune* story (Webster & Corey, 2021).

Metropolitan Urban Service Area, which is served by the regional sewer system. This area includes almost all of the incorporated cities in the seven-county metro area.

Most cities, or their contracted vendors, were able to provide parcel-level shapefiles soon after we filed requests in summer 2019. Some shapefiles showed only the locations of the zoning districts. We turned these into parcel-level files by doing a spatial join between the district shapefile and a parcel shapefile that county property record offices jointly make available to the public (Minnesota Geospatial Commons, 2022).

We loaded each city's parcel shapefile into a PostGIS database to facilitate spatial operations more efficiently. The resulting parcel database shapefile also provided data points, such as estimated value and year built, for most properties.

To standardize the zoning districts, we categorized each one by whether it was residential or nonresidential and then further categorized the residential districts by the densest type of housing allowed by right in that district, according to the district definitions provided in each city's zoning ordinances. The types of districts follow:

- Single-family zoning allows only detached houses.
- "Middle housing" zoning allows small, low-rise two- to four-family buildings, mobile homes, or townhouses in addition to single-family dwellings.
- Multifamily zoning includes a wide range of residential densities. Multifamily zoning, in our taxonomy, includes mixed-use zones that allow multifamily housing in addition to commercial uses.
- Planned Unit Development zoning is our final category. A large proportion of recent growth in the Twin Cities region has taken place in planned developments (*Rights, responsibilities, and preemption in Minnesota*, 2022), which are

regulatory limits customized to a specific project (Mandelkar, 2017). In some cases, as in the city of Minneapolis, the city makes PUD agreements but officially leaves a standard zoning designation in place. In other cases, especially when a PUD is used to develop previously agricultural land, the land is zoned “PUD” or similar and has no zoning regulation beyond the custom PUD agreement.

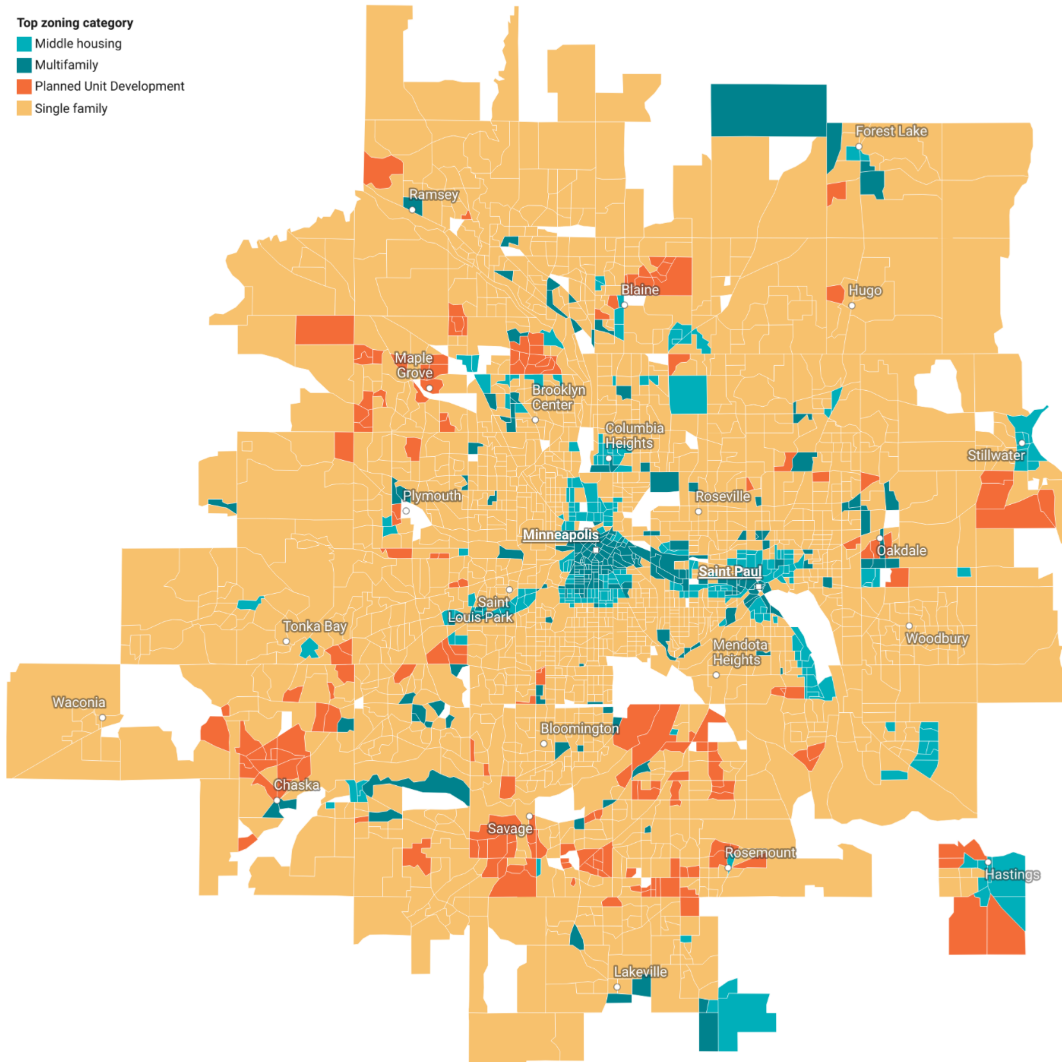
For each block group, we calculated the share of residential land zoned for single-family, middle, and multifamily housing. For some records, but not all, zoning data identified the actual lot size, house size, and appraised value. Separately, we measured the share of nonresidential zoning as a share of all zoned land in the block group. As Figure 3 shows, single-family zoning is the predominant category in the metro area. We measured each type of residential zoning listed as a share of residentially zoned land in the block group.

In one substantial case, we were uncertain how to code a zoning district. Bloomington’s R-1 zone, which covers most of that city’s residential land, allows duplexes, but only under strict conditions, including a large lot size and deeper setbacks than are allowed for a single-family house. We found that, unlike in other middle housing zones, middle housing was rare in Bloomington’s. Thus, we coded the zone as single family only. For robustness, we also ran the analysis with the R-1 zone coded as middle housing; it made little difference.

We spatially joined the zoning data to census block group data from the 2020 Census and 2020 and 2019 five-year American Community Survey (ACS).

Figure 3: Residential Zoning in the Twin Cities

Which type of residential zoning is most common in each block group? Note that most block groups include multiple types of zoning.



Map: Salim Furth and MaryJo Webster • Created with Datawrapper

We used data from the 2020 Census to measure the racial and ethnic composition of each block group. And we used the ACS to measure housing age and median gross rent. In some block groups, median housing age was unavailable, so we estimated it from the data of nearby block groups using spatial smoothing. Median gross rent was available only at the census tract level.

For household-level empirics, such as the analysis of immigrant presence by race displayed in Figure 2, we used ACS data published by Ruggles et al. (2021).

Table 1 shows that housing typology closely follows zoning designations in block groups zoned almost exclusively for single-family or multifamily housing. Half of the units in middle housing zones, however, are single-family detached houses, and just a third are middle housing typologies. PUD block groups display a broad mix of typologies.

Table 1: Housing Typology Follows Zoning

	Housing units reported (%)			Number of block groups
	Detached	Middle	Multifamily	
Zoned single-family detached	88	7	5	510
Zoned middle housing	53	35	12	29
Zoned multifamily	3	11	85	97
Zoned Planned Unit Development	45	31	24	45

Sources: Municipal zoning data; American Community Survey (ACS) 2020 5-year data.

Note: This table shows the housing typologies reported by ACS respondents in block groups where residential zoning falls overwhelmingly (over 97%) into a single category. Block groups may contain parcels zoned for nonresidential use and multiple zoning districts within the same housing type category. Middle housing includes townhomes, mobile homes, and two- to four-unit buildings. The data do not reflect Minneapolis's 2020 upzoning.

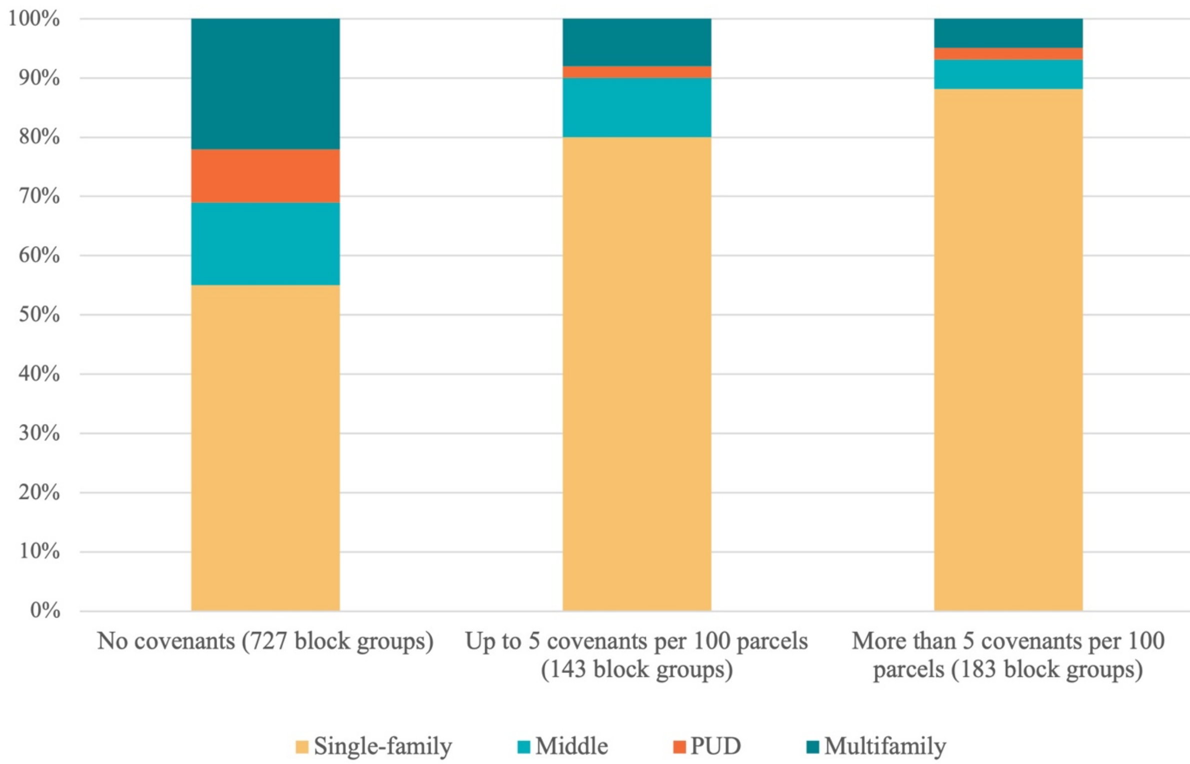
Following Sood and Ehrman-Solberg (2022), we used the database of Hennepin County racial deed restrictions published by the Mapping Prejudice Project at the University of Minnesota (Ehrman-Solberg et al., 2020). We cleaned the data by removing 3,864 deed restrictions on cemetery plots and duplicates. Then we assigned each covenant to a 2020 block group and divided the number of covenants by the number of present-day residential parcels. In two block groups, there were more covenants than parcels, likely because covenanted houses were removed for highway construction.

Although we have not matched zoning to covenant data on a parcel-to-parcel basis, it is clear that the overwhelming majority of covenants in Hennepin County were associated with single-family houses.[12]

[12] We identified one exception in the data: 11 individual apartments in the building at 501 Theodore Wirth Parkway had deed restrictions.

As Figure 4 shows, single-family zoning covers just under half of land in the 69% of Hennepin County block groups where no covenant was recorded, but 89% of land in block groups where covenants were common. It also shows that a large number of block groups had just a few covenants.

Figure 4: Residential Zoning by Prevalence of Racial Covenants



Data sources: Mapping Prejudice Project Hennepin County racial covenants data; Ehrman-Solberg et al. (2020); municipal zoning data.

Note: This figure shows the distribution of residential zoning types in Hennepin County block groups with varying concentrations of racial covenants. Racial covenants typically prohibited occupancy by "non-Caucasians" and were deemed unenforceable in 1948. Most covenants were recorded in the 1930s and 1940s; a few were recorded as late as 1956. Residential parcel counts and zoning were measured in 2019. Because of the time lag between the count of covenants and the count of residential parcels, the ratio between covenants and parcels exceeds 1 in a few block groups where demolitions occurred. The data are not weighted. The data do not reflect Minneapolis's 2020 upzoning.

5. Methods

Demographics, zoning, the built environment, and ownership patterns are co-determined in a web of two-way causal relationships that are, under normal circumstances, impossible to untangle. To firmly establish the strength of any one causal link in such an environment, a researcher normally needs a "natural experiment" or similar mechanism that introduces changes on an effectively random basis. Planners, builders, and residents, however, strive to optimize their choices, thoughtfully avoiding randomness.

A less ideal, but plausible case for causation can be made when one decision clearly precedes another. For instance, zoning districts created a generation ago can be reasonably viewed as causes of current construction and residential choices, as in Twinam (2018). However, even though most zoning districts remain little changed over decades, the incidence of rezoning is decidedly nonrandom (Whittemore, 2019).

Resseger (2022) and Kulka (2021) argue that their estimates are causal because they are based on proximity: they consider bordering census blocks and tracts, respectively, and argue that those are so similar in other regards that the only difference between them is zoning. One can certainly imagine exceptions. For instance, if multifamily housing is often used to buffer single-family housing from nuisances, multifamily blocks will systematically differ from single-family blocks in distance from nuisances as well as zoning.

Finally, even where research yields a valid causal estimate of the impact of longstanding zoning designations on demographics (or typology, prices, or traffic), its external validity—and thus value to policymakers—is limited. When land is first converted to urban use, the prevailing zoning immensely influences what is built there. But the pace and depth of change following a broad rezoning are highly dependent on existing conditions and prevailing prices.

In this fraught environment, why attempt to link zoning and race at all? Because, unlike construction patterns, prices, ownership, and migration, zoning is a direct policy lever.[13] Even in the absence of clear causal estimates, cities must make zoning choices, and they ought to have as much information as research can afford.

Given the difficulty of identifying causal, externally valid estimates, we prefer to think of most estimates, including ours, as increasingly accurate descriptions of the prevailing relationship between zoning and race.

Empirical Approach

To construct our description of that relationship, we regress the non-White population share on multifamily, middle housing, PUD zoning, minimum lot size, and nonresidential zoning with spatially correlated errors and geographic control variables:

NON-WHITE = f(MULTI, MIDDLE, PUD, NONRESIDENTIAL, GEOGRAPHIC CONTROLS) + ERROR

Our residential zoning categories are expressed as a percentage of each block group's residential land. The omitted category is single-family zoning, so the coefficients on the other residential zoning shares can be interpreted as the increase in the non-White population share associated with a higher land share in the respective zoning category at the expense of single-family zoning.

Nonresidential zoning is measured as a percentage of all zoned land in the block group. Our geographic controls include an indicator for block groups closer to Minneapolis's downtown

[13] To be sure, governments at various levels influence construction patterns, prices, ownership, and migration. But those indirect and often uncertain interventions cannot be reasonably compared to zoning, which is city created and city enforced.

than to St. Paul's and the linear distance from the nearest downtown. We include an interaction term for the last two variables, which helps reflect the fact that the St. Paul side of the metro area is smaller. Our regressions include either fixed effects for every city and town or an indicator variable for the central cities. We found that racial population shares have a nonmonotonic relationship with the age of the housing stock, so we use polynomials of median housing age up to a quadratic term.

In some specifications, we include minimum lot size or the prevalence of old racial covenants, both discussed in the following sections. Other occasional controls include median house size and housing cost variables.

Spatial regression methods are essential for accurately estimating geographic patterns that are likely influenced by myriad factors that change gradually across space and that spill over boundaries (Kelly, 2019). Our preferred approach is generalized spatial two-stage least squares (GS2SLS), which iteratively estimates the spatial correlation of errors. We chose to use a truncated inverse distance error correlation matrix. This embodies our assumption that the correlation between two places declines geometrically with distance, and disappears at a sufficient distance.[14] Throughout, we use the `spregress` command in Stata.

It is also possible to include spatial lags of the dependent and independent variables in a spatial regression, allowing spillovers from each block group to those nearby. Although perhaps attractive, further spatial interdependence adds interpretive difficulty without addressing the question at hand.

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[14] Truncation is important because the outer ring of the metro, in all directions, is heavily White. Thus, as an artifact of limiting our sample to a roughly circular metro area, there is a very high correlation in the White population share distances approaching the region's diameter. Truncating fixes this problem.

Gibbons and Overman (2012) argue that spatial interdependence should be included only when a clear theoretical or contextual reason justifies it. Practically, spatial interdependence also introduces greater dependence on software. In our data, for example, we find radically different results when we regress the White population share on its own spatial lags and some set of independent variables than when we do the same exercise using the non-White population share as the dependent variable.

To decompose the impact of zoning across differing racial and ethnic groups, we regress each major group's population share on zoning independently. We also investigate whether zoning's relationship with race differs in the suburbs and cities.

To benchmark our preferred GS2SLS results, we also perform ordinary least squares (OLS) regressions with standard errors clustered at the city level.

Minimum Lot Size

The only density restriction that we could compare across all jurisdictions is minimum lot size, which is a limiting constraint for single-family and duplex development. However, we expect that smaller lot size minimums are likely correlated with more liberal limits on multifamily development, which would confound our results.[15]

Thus, to test the impact of minimum lot size on racial demographics, we limit our sample to places that are overwhelmingly single-family zoned. We verified that recently built houses largely follow published minimum lot sizes. Just 22% of houses built since 2000 clearly deviate, being sited on lots at least 10% smaller than the published (c. 2020) minimum. Another 51% of 21st-century houses are on lots between 90% and 150% of the published minimum. Research on

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[15] Surprisingly, there is very little correlation (−0.10) between median minimum lot size and the shares of residential land devoted to multifamily and middle housing.

growing cities in Texas has found higher rates of deviation (Gray & Furth, 2019; Gray & Millsap, 2020).

We prefer to measure minimum lot sizes in logarithmic terms, which follows a conceivable utility function in land: as lot size increases, willingness to pay for an additional square foot gradually falls. An alternative, used by Kulka (2021) among others, is to use dwelling units per acre (DUPAC), which implies that willingness to pay declines linearly in the density of neighbors. Relative to our preferred specification, DUPAC magnifies differences between small lots and minimizes differences between large lots. We test both specifications and also test the effect of large-lot outliers.

Racial Covenants

Given that racial covenants are strongly associated with single-family development, as Figure 4 shows, it is possible that the relationship between zoning and race is in part an artifact of the long shadow of the covenants. Alternately, it is possible that the relationship Sood and Ehrman-Solberg (2022) observe between covenants and current racial makeup is in part a consequence of zoning.

We first verify that the relationship between zoning and covenants is not merely a geographic coincidence. Even when controlling for city, distance to downtown, and housing age, we see that covenants are a strong predictor of a block group's single-family zoning share.

We then loosely follow Sood and Ehrman-Solberg's setup, adjusted to our own demographic data and geographic controls. Because of the distinct doughnut pattern of racial covenants, we include a term for squared distance from downtown.

Finally, we include zoning variables along with covenant prevalence to measure the relative strength of their relationships to non-White population share.

6. Results

In this section, we preview our main results and then walk through the battery of regressions outlined in the previous section. Table 2 provides a summary of the key results from our preferred specification. A block group zoned entirely for multifamily housing has a 21-percentage-point larger non-White population share than an equally situated single-family block group. Middle housing and PUD zones are associated with non-White population shares 14 and 8 percentage points, respectively, larger than single-family zones.

Table 2: Key Results Summary

If a block group's residential land is entirely zoned for …	Multifamily	Middle housing	Planned Unit Development
… how much higher, in percentage points, is each group's population share than if it were zoned single-family?			
Non-White (32%):	21	14	8
Black (11%):	18	6	6
Asian (8%):	−1	0	2
Hispanic (7%):	4	7	0
Others (6%):	4	8	0

Note: This table summarizes the results of Regressions (1) and (5)–(7). The full results, including standard errors and control variables, are reported in Appendix Tables A.1 and A.2. The relationship between zoning and race is reported here in percentage points of the population. These are not causal estimates. Each row is independently estimated, and errors are presumably correlated across regressions, so columns do not add up. Each race's metro area share is included in parentheses. All nonzero results in this table are statistically different from zero. Ethnic Hispanics are not included as members of any other race. "Others" consists mostly of multiracial people, in addition to Native Americans and Pacific Islanders.

Table 2 also reports the key implications of Regressions (5)–(8), reported in Appendix Table A.2, which duplicate the analysis for each large racial or ethnic minority. The effects of multifamily and PUD zoning are concentrated among Black households. If zoning affected each non-White group equally, we would expect the coefficients to be in proportion to that group's share of the metro area population, which is noted in parentheses in Table 2. Instead, about three-fourths of the effects of multifamily and PUD zoning are concentrated in the one-third of the non-White population that is Black.

Zoning for middle housing, by contrast, has relatively stronger effects among Hispanic and other households. We find almost no difference between Asian and White residency patterns.

Key Results

Our preferred specification is given in Regression (1), in Appendix Table A.1. In this and all following tables, we inflate the dependent variable by a factor of 100 to make the coefficients easy to read. All of our zoning metrics and geographic controls are statistically significant. Aside from the key results, previewed previously, we find that nonresidential zoning is associated with a higher non-White population share.

The non-White population declines with distance from the center in areas nearer downtown. The decline is twice as rapid on the St. Paul side of the metro. Block groups in the center city have non-White populations 10 percentage points higher, all else equal. Housing age, which we express using four polynomials, has a nonmonotonic but statistically significant relationship with race. None of the geographic coefficients, however, is as robust as the zoning coefficients. As we change around the set of geographic controls and sample size, individual effects drop in and out of significance and occasionally change sign.

In Regression (2), we replace several of the geographic controls with municipality indicators. The coefficients of interest are unchanged.

In Regression (3), we introduce more controls, including the median minimum lot size and house size for detached houses in each block group. Doing so reduces our sample, because house size data were unavailable for most of Hennepin County, but it shows that the headline results are robust to the inclusion of local housing attributes. Minimum lot size has a negative and significant relationship with non-White population share, but the magnitude is modest.

In Regression (4), we add estimates of median home value (by block group, from tax appraisals) and median gross rent (by census tract, from the ACS). We would not be discouraged if the inclusion of price variables pushed the zoning coefficients to statistical insignificance, because zoning may influence demographics purely by raising housing costs. However, the inclusion of price and town variables has no statistical impact on our core results. This suggests that zoning affects racial demographics through channels other than prices.

In Regressions (3) and (4), the coefficient on middle zoning is smaller than in Regression (1). Otherwise, the coefficients of interest do not differ significantly across the regressions in Appendix Table A.1.

In Appendix Table A.2, we reproduce Regression (1) for the Black, Asian, and Hispanic population shares separately, and for "other" residents who do not fit into any major group.[16] It would be convenient if the rows of Appendix Table A.2 summed neatly to the coefficients of Regression (1). However, each regression is independently performed, and the results do not sum, perhaps because error terms are correlated across regressions. As noted, we found disproportionately strong effects of multifamily and PUD zoning for Black residents and stronger effects of middle zoning for Hispanic and other residents. Asian population shares are almost completely uncorrelated with zoning.

In Appendix Table A.3, we repeat Regressions (1) through (4) using OLS, a standard statistical benchmark. To account for some spatial correlation, we cluster standard errors at the city level. The OLS results are less precise, reflecting the inferior handling of spatial correlation. In Regressions (9) and (10), which are parallel to (1) and (2) in Appendix Table A.1, the

[16] "Other" consists of mostly mixed-race people.

coefficient on middle zoning is larger than that on multifamily zoning, although the difference is not statistically significant.

To verify that our results are robust, we repeat Regression (1), with slight variations, for three subsamples, which we report in Appendix Table A.4. Regression (13) includes only block groups where the median housing unit is less than 35 years old. Although this regression includes only 30% of the sample, its key coefficients barely differ from those in Regression (1).

Regressions (14) and (15), respectively, cover the suburbs and cities alone. PUD zoning barely exists in the cities, so we omit it. The association of multifamily zoning with race is stronger in the suburbs than in the full sample. The rest of the coefficients are not significantly different from the baseline estimates.

Minimum Lot Size

We find that minimum lot sizes have a less clear relationship with racial patterns. The size and statistical significance of the association depends on how we define density (log lot size or units per acre) and whether we include areas zoned for very large lots. Appendix Table A.5 gives detailed regression results, and Table 3 summarizes their effects. A difference of 5,000 square feet in minimum lot size can be associated with as much as 3.9 percentage points or as little as 0.4 percentage points difference in non-White population share.

Table 3: Measuring Minimum Lot Size Differences

Compared to a block group with a 10,000 square foot lot minimum lot size, how does the non-White population share differ, in percentage points, in a block group where the minimum is …			
		5,000 sq ft	15,000 sq ft
According to …	Regression (16):	0.8 higher	0.4 lower
	Regression (17):	3.9 higher	1.3 lower
	Regression (18):	3.7 higher	2.1 lower

Note: This table summarizes the relationship between minimum lot size and race, as estimated in Regressions (16)–(18), in comparable terms. See Appendix Table A.5 for full results, including standard errors and controls. The first numerical column presents the average difference in the non-White share of the population between a single-family block group where the median minimum lot size is 5,000 square feet (sq ft) and one where the median minimum lot size is 10,000 sq ft, controlling for geography. The last column compares a block group zoned for 15,000 sq ft lots to one zoned for 10,000 sq ft lots. The three lot sizes represent the 10th, 50th, and 91st percentiles, respectively, of median minimum lot size in the 2,183-block group sample for which data are available and the 10th, 50th, and 87th percentiles, respectively, in the 479-block group subsample of areas zoned at least 97% single family. The coefficient for Regression (16) is not statistically different than zero.

In our preferred specification, Regression (16), we find a small, statistically insignificant association between minimum lot sizes (expressed in logarithmic terms) and the non-White population share. But in Regression (17) we express minimum lot size in units per acre, following Kulka (2021), and find a strong effect. We find that a four-unit-per-acre difference in allowed density is associated with a 3.6-percentage-point difference in non-White share, which is statistically indistinguishable from Kulka's estimate of 5 percentage points.

A likely cause of the discrepancy between using log minimum lot size and DUPAC is the way in which the two approaches treat outliers. The units-per-acre approach minimizes differences among large-lot zones, giving them very little statistical leverage. But in our preferred logarithmic specification, they have ample leverage.

In Regression (18), we repeat our preferred approach, this time excluding large-lot zones (areas with median minimum lot sizes of 2.5 to 10 acres).[17] In contrast to the small, weak

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[17] An additional reason to exclude these zones is that a large share of their house lots are subdivided to conventional suburban densities, far below the official minimum lot size.

coefficient in Regression (16), the censored sample yields a strong coefficient, comparable to the one yielded by the unit-per-acre approach.

The power of outliers arises in part because there is little variation in the rest of the sample; fully 60% of tracts in the subsample have lot sizes between 10,000 and 12,000 square feet. This sensitivity and lack of variation suggest that our data sample is not ideal for investigating the effects of minimum lot sizes.

In unreported regressions, we found that the results are robust to controlling for median house size, but doing so cuts the sample by almost half. Including county indicators pushes the key coefficient in Regression (17) into statistical insignificance.[18] The results are all slightly stronger when we limit the sample to suburbs.

Interestingly, controlling for median home price completely eliminates the association between race and minimum lot size, however expressed. We show this in Regression (19). This result stands in contrast to Regression (6), which found that home prices and rent, although important predictors of racial demographics in their own right, did not significantly attenuate the relationship between typology zoning and race.

To ease the comparisons between the coefficients, Table 3 summarizes the associations between lot size and race implied by Regressions (16) through (18). The coefficient for Regression (18) implies that zoning for 5,000-square-foot lots rather than 10,000-square-foot lots is comparable to zoning 18% more residential land for multifamily at the expense of single family.

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[18] We cannot reasonably include municipality indicators because most municipalities have no variation in median minimum lot size.

Racial Covenants

The coincidence of pre-1956 racial covenants and single-family zoning, shown in Figure 4, implies that studying either one in the other's absence might introduce omitted variable bias. However, our joint analysis shows that both retain explanatory power.

Because we use the same data on covenants as Sood and Ehrman-Solberg (2022) (although other data sources differ), we follow them in separately analyzing the effects of covenants for all non-Whites (in Appendix Table A.6, which includes Regressions [20] through [24]) and for Black residents only (in Appendix Table A.7 and Regressions [25]–[29]).

We begin our analysis with Regressions (20) and (25), which are loosely analogous to Sood and Ehrman-Solberg's setup: the sample is limited to the city of Minneapolis, and covenants are expressed as the hyperbolic inverse sine of the ratio between recorded covenants and modern residential parcels.[19] We find a larger and statistically significant effect for the non-White population share but not the Black share. Sood and Ehrman-Solberg find the opposite (2022, p. 31).

As we show in Figure 4, many block groups have a very small number of covenants. We look for a threshold effect by including an indicator variable that equals 1 if a single covenant is recorded in a block group and 0 otherwise. In Minneapolis proper (Regressions [21] and [26]), we find no evidence of a threshold effect. But when we extend the sample to cover all of Hennepin County, in Regressions (22) and (25), the threshold effect becomes important, and the intensity effect fades.

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[19] Although this function is concave, the concavity over the sample range [0,1.85] is extremely small, so that the hyperbolic inverse sine can be almost perfectly approximated with a linear function.

The key regressions in this section are (23) and (28), which show that typology zoning remains a powerful predictor of non-White and Black population share, respectively, in the presence of racial covenant data. The threshold effect for covenants drops by about half, and the intensity effect fades further into statistical insignificance.

Finally, in Regressions (24) and (29), we provide a benchmark for the zoning coefficients, which prove to be statistically unchanged by the presence of covenant data. We find it interesting that the explanatory power of the countywide regressions is about equal whether we use covenant data alone, zoning data alone, or the two together, along with a common set of geographic controls.[20]

7. Discussion

How, exactly, do restrictions on building types result in racial sorting? Following Pendall (2000), we hypothesize a "chain of exclusion" that links typology zoning to race. In this section, we briefly document evidence for each link in the chain, referring to existing research and, where possible, verifying that national patterns hold in our data as well.

We hypothesize a simple, linear chain of exclusion, without denying that many secondary avenues exist:

- Zoning determines which housing typologies predominate in an area.
- Housing typologies have differing ownership rates.
- Ownership rates differ sharply by race.

[20] The geographic controls by themselves have somewhat less explanatory power. Regressing non-White and Black population share on the controls alone yields pseudo R^2 of 0.27 and 0.24, respectively.

Along the way, this section addresses two side issues: Does the preponderance of immigrants among the Twin Cities' non-White population affect the chain of exclusion? And do cultural preferences for some typologies help explain the varying relative importance of multifamily versus middle housing among race and ethnic groups?

Zoning Determines Typology

Implicit in our empirical strategy is the belief that zoning is a meaningful summary of the possibilities for local housing. However, there are a number of reasons that built typology may depart from zoning.

Districts include some nonconforming units that predate the current zoning, were created illegally (e.g., by subdividing a house into a duplex), or were allowed within a Planned Unit Development that is not reflected in the published zoning.

We have coded zones by the densest typology allowed. But residential districts are typically additive, so that single-family homes are allowed in multifamily and middle-housing districts. Depending on market conditions, a district may be built up much less than its zoned potential. Recently upzoned areas can easily remain less dense than their new designation for decades.

Despite these possibilities, Table 1 shows that multifamily and single-family detached typologies dominate multifamily and single-family zones, respectively. Middle housing zones include a large number of single-family detached houses, introducing a weak link in this chain. We would presumably find stronger effects of middle zoning if it were more tightly tied to typology.

Different Typology, Different Ownership

Multifamily buildings are mostly rented, while single-family dwellings are mostly owner occupied. Table 4 shows occupant ownership rates across a range of typologies. Ownership is higher in the Twin Cities seven-county metro area than nationally for single-family and small multifamily buildings. However, the Twin Cities area has slightly lower ownership for units in midsized and large multifamily buildings.

Table 4: Owner Occupancy Rates by Typology

	Owner occupancy rate (%)	
	National	Twin Cities
Single-family detached	84	93
Townhouse	62	70
Duplex	23	26
Triplex or fourplex	13	22
5- to 9-family building	11	20
10- to 19-family building	9	7
20- to 49-family building	12	12
50-family building and up	15	14

Sources: American Community Survey (ACS) 2019 5-year data; Ruggles et al. (2021).

Note: This table shows the differing rates of owner occupancy for housing typologies distinguished in the ACS. The Twin Cities data cover the seven-county metro.

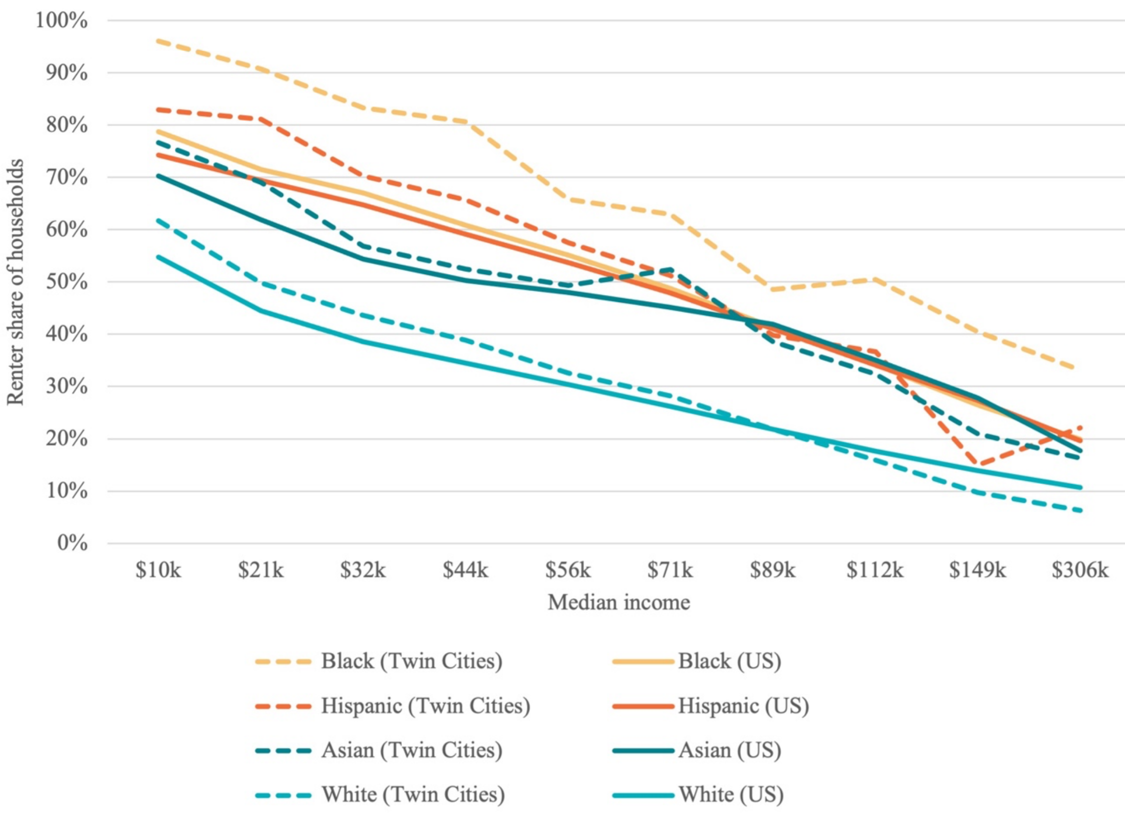
There are three obvious explanations for the drastic difference in ownership rates between single- and multifamily dwellings. First, single-family houses cost more and ownership is more attainable for wealthier people. Second, the benefits of ownership, such as designing and landscaping to one's own tastes, are more easily achieved on a detached house lot. Third, the efficiencies of managing many rental units are more easily achieved in a multifamily setting.[21]

[21] The emergence of large single-family rental subdivisions, especially in the U.S. South, may weaken the relationship between race and typology zoning. Using data from nine Florida counties, Ihlanfeldt and Yang (2021) find that a 2-percentage-point increase in single-family rentals leads to a 1-percentage-point increase in Black population, offset by an equal decrease in Hispanic population.

Ownership Rates Differ Sharply by Race

Non-White Twin Cities households are much more likely than White households to rent their homes. Part of the explanation is the difference in incomes. However, as Figure 5 shows, non-White households are much more likely than White households to rent at all income levels. The gap between White and non-White households at the same income level helps explain our finding in Regression (4) that home prices and rents do very little to explain the relationship between multifamily zoning and race.

Figure 5: Rentership by Race and Income



Sources: American Community Survey (ACS) 2019 5-year data; Ruggles et al. (2021).

Note: This figure shows the share of households headed by a person of each race or ethnic group that rent. We calculate the share within each decile of the national, all-race income distribution. Households that do not rent own their dwelling. Rates for the Twin Cities metro area are imprecise owing to small sample sizes. The x-axis shows the mean national income within the decile.

As Choi et al. (2019), among others, have noted, the Black–White homeownership gap is especially wide in the Twin Cities. Using ACS data from Ruggles et al. (2021), we find that Black-headed households earning around $100,000 per year have a homeownership rate about equal to White-headed households earning around $20,000.

Is this enormous gap the result of differing observable household characteristics? One initially appealing explanation for the large Black–White homeownership gap is the large share of Black Minnesotans from recent immigrant communities. However, a quick look at the data shows this explanation is unlikely. Black immigrant households in the Twin Cities have higher household incomes and a higher homeownership rate than African American households.

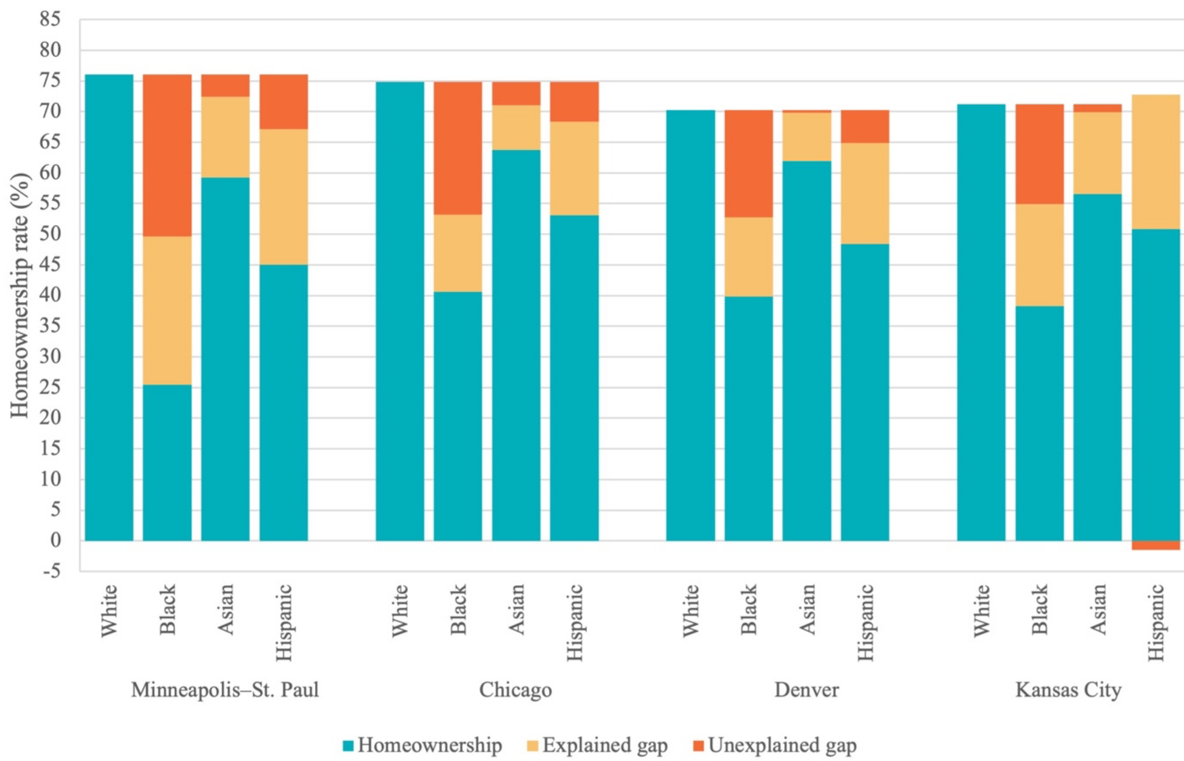
More formally, in a logit model of homeownership (unreported), we find no evidence of a difference in homeownership between Black households with and without an immigrant member. (Nor do we find such a difference for other races.)

Looking further for potential explanations for the extreme Black–White homeownership gap in the Twin Cities area, we compare the 35 largest U.S. metro areas using metro-specific Blinder–Oaxaca decompositions.

A Blinder–Oaxaca decomposition splits the racial gap into a part that can be explained by differences in observable characteristics, such as income and age (see Appendix Table A.8 for a complete list), and a part that is a remaining, unexplained gap (Sinning et al., 2008). In the variant we employ, the software performs a logistic regression of homeownership status for White-headed households in each metro. It then predicts homeownership for each non-White household using the coefficients from the logistic regression. In a few cases, the predicted homeownership rate for a minority is lower than the actual homeownership rate—this is the case for Hispanics in Kansas City, for instance. But in most cases, the predicted homeownership rate is higher, leaving an unexplained gap. It is likely that a substantial part of the unexplained gap would be explained if we had data on credit rating factors and family wealth (Dey & Brown, 2022; Brandsaas, 2021).

Figure 6 shows the results of the decomposition for the Twin Cities and three other metro areas. Appendix Table A.9 shows results for all 35 metros. Minneapolis–St. Paul has both the largest explained gap and the largest unexplained gap between Black and White homeownership. As in most metros, the Twin Cities' unexplained gap between White and Hispanic homeownership is much smaller.

Figure 6: Homeownership Gaps by Race in Selected Metro Areas



Sources: American Community Survey (ACS) 2019 5-year data; Ruggles et al. (2021).

Note: This figure compares observed homeownership rates for minority-headed households to those of White-headed households in four metropolitan areas. The teal column base shows the actual homeownership rate. For each minority in each metro area, we used a Blinder–Oaxaca decomposition to identify the portion of the minority–White gap that can be explained by the observable characteristics listed in Appendix Table A.8. The explained portion is shown in yellow. We lack data on credit characteristics and parental wealth, which would explain still more of the gap. The unexplained portion of the gap is in orange. In the case of Hispanic households in Kansas City, observable data slightly overexplain the gap, leaving a small negative unexplained portion. See Appendix Table A.9 for results across 35 metro areas.

The relative size of the homeownership gaps, both raw and unexplained, correspond to the relationships between zoning and race that we found previously: largest for Black households, smallest for Asian ones. This finding is consistent with our hypothesis that the association between zoning and race is mediated by differences in ownership. Thus, we have evidence for every link in the chain of exclusion: zoning is closely linked to typology, typology is closely linked to tenure, and tenure differs sharply across races.

Differing Typology Choices across Racial and Ethnic Groups

One alternative to our hypothesis is if similar renter or owner households of different races chose different housing typologies because of discrimination or differences in cultural norms. Such discrimination would violate Minnesota law even when it is legal under the federal Fair Housing Act, which does not cover single-family houses and small, owner-occupied multifamily buildings.[22]

Our core findings also include differences between similarly situated minorities that are not explained by the simple chain of exclusion. Recall from Table 2 that middle housing and PUD zoning have a disproportionate association with Hispanic and other residents, whereas multifamily zoning has a disproportionate association with Black residents.

We use ACS microdata from Ruggles et al. (2021) to quantify differing typology choices by race in two ways. First, we run logistic regressions predicting housing typology depending on race and other observables.

[22] Minnesota law significantly narrows the Fair Housing Act's exemptions. Racial discrimination is not allowed even for room rentals within an owner-occupied home (Human Rights, 2021; U.S. Department of Housing and Urban Development, n.d.).

For this exercise, we narrow our consideration to households living in detached houses or middle and multifamily buildings of three or more units. And we include only households in public use microdata areas (PUMAs), where single-family and three-plus unit buildings each account for at least 30% of occupied dwellings across those two categories, so households have ample choice. Appendix Table A.10 reports the results.

Using data from the seven-county Twin Cities metro, we find no significant difference in typology preference by race.[23] We do find one difference across immigrant status, in Regression (33): Black renter households with an immigrant member were significantly more likely to occupy multifamily dwellings than otherwise similar Black, nonimmigrant households in the same PUMA.

Our second exercise, shown in Table 5, is to compare the rates of single-family, middle housing, and multifamily occupancy for renter households of six racial and ethnic groups without controlling for other differences, such as income and child presence. Because the previous exercise found a significant difference between immigrant and nonimmigrant Black households, we distinguish "Black immigrant" from "African American" households in Table 5 on the basis of immigrant presence.

[23] Using national data in Regressions (30) and (31), we find that Black owners favor detached houses over multifamily buildings more so than White owners. Investigating this regularity geographically in unreported metro-level regressions, we find that it is driven by a strong proclivity for detached homes among Black owners in Southern metro areas, such as Memphis and Orlando. National data also show that White renters favor detached houses relative to Black, Hispanic, and Asian renters.

Table 5: Renters' Residences in the Twin Cities Metro

	Share of housing (%)					
Racial or ethnic group	Single-family detached	Middle housing	Multifamily	Median income ($)	Child presence (%)	Number
White	14	20	66	49,941	16	229,066
African American	13	20	66	30,215	36	39,750
Black immigrant	6	21	72	32,682	50	33,646
Asian	12	18	70	55,079	43	26,358
Hispanic	14	27	58	40,466	54	27,800
Other	14	29	57	34,531	35	15,575
All	13	21	66	44,243	27	372,195

Sources: American Community Survey (ACS) 2019 5-year data; Ruggles et al. (2021).

Note: This table records the residential typologies of Twin Cities renters by race and ethnicity of the head of household. We divide housing into single-family detached, middle housing, and multifamily housing. Middle housing includes townhomes, mobile homes, and two- to four-unit buildings. Multifamily refers to buildings with at least five units. Following the finding in Regression (33) of differing residential choices among Black households depending on immigrant presence, we divide Black households into those with and without an immigrant member. The latter we label "African American." We also record median income and the percentage of renter households with a child present, which regressions showed to be significant predictors of housing typology choice. The number of households is the sum of weights in the ACS.

The differences between the six groups are muted. White and African American renters have indistinguishable typology patterns. Black immigrant households are about half as likely as others to rent detached houses. Hispanic and other households are substantially more likely than other groups to rent middle housing and less likely to rent multifamily units. This finding corresponds to our finding that the Hispanic and other population is correlated with middle zoning.

Differences across groups could be driven by geographic choices—for instance, if Hispanic families happen to concentrate in areas with high shares of middle housing. To test for this possibility, we construct counterfactual renter shares for each group that are based on the breakdown of the housing stock by PUMA. This measure is almost identical across all six groups, implying that the variation is not caused by geographic differences. Instead, almost all the variation reflects different choices made by people facing roughly the same choice set.

8. Conclusion and Policy Implications

This paper has added new empirical evidence to the large qualitative body of work showing that zoning has the effect of sorting racial minorities into relatively small districts zoned for multifamily and middle housing. In the Twin Cities metro, which has unique ethnic demographics, we find strong, robust associations between zoning and racial residential patterns. An area zoned for multifamily housing has, on average, 21 percentage points more non-White residents than an identically situated single-family zone.

A district that allows fourplexes, duplexes, townhouses, or mobile homes in addition to detached houses has, on average, 14 percentage points more non-White residents than a single-family district, although detached single-family houses remain the most common dwelling type in middle housing zones. Hispanic and mixed-race people appear especially likely to live in middle housing zones.

In the growing suburbs of the Twin Cities, many towns rely on discretionary PUD zones. As a group, PUD zones have a mixed housing stock. Not surprisingly, we find that they have significantly higher non-White population shares than single-family zones.

Our work is fundamentally descriptive. We hope that policymakers will find it useful as they seek to understand the exclusionary history and effects of zoning and America's persistent failure to achieve racial integration.

However, the relationship between zoning and race—even from studies with strong causal identification—is bound to a time and place. Zoning, as we study it, is an expression of many overlapping institutions and decisions over time.

Shortly after our study period, Minneapolis made national headlines by upzoning its single-family neighborhoods to allow duplexes and triplexes. St. Paul policymakers may follow (Yudhishthu, 2022). However, Minneapolis's legislative change is not rapidly transforming either the housing stock or the demographics of the affected neighborhood.[24] We applaud the reform—but we expect its transformation to be wrought over decades, not years.

Exurban, greenfield zoning would likely come closer to mirroring our findings—zoning for denser typologies would likely yield more rental units and thus a more representative population. But our study was not limited to recently developed sites, and it is entirely possible that narrower estimates would show a weaker or stronger relationship.

Although we must remain agnostic on the magnitudes of demographic change that would accompany reforms in different contexts, the clear links from zoning to typology to tenure to race give policymakers a clear framework for thinking about zoning and race.

As the United States confronts its history of systematic public and private housing discrimination, zoning is one of the prime targets for reform. Our evidence shows that zoning exclusively for detached houses effectively excludes large numbers of non-White households—as well as many white ones—who rent. It is especially important to ensure that multifamily and middle housing zoning are allowed in newly urbanizing areas where zoning has its greatest effect. At the same time, even drastic zoning deregulation would partially address just one of many remaining racial disparities. Gaps in family wealth, marriage, educational opportunities, income, and homeownership are large and upstream of residential possibilities and choices.

[24] Duplex and triplex construction in Minneapolis has been especially slow because the city kept regulations, including height, side yard, and floor area ratios, that effectively prevent the creation of duplexes or triplexes with large units (Pinto et al., 2022).

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Appendix

Appendix Table A.1: Generalized Spatial Two-Stage Least Squares Results
Dependent variable: Non-White population share × 100

Regression	(1)	(2)	(3)	(4)
Multifamily zoning	20.8***	20.3***	20.9***	20.7***
	(1.4)	(1.4)	(1.8)	(1.4)
Middle zoning	14.0***	14.0***	8.1***	8.8***
	(1.3)	(1.3)	(1.4)	(1.2)
Planned Unit Development zoning	7.7***	7.9***	8.0***	6.6***
	(1.4)	(1.4)	(1.6)	(1.5)
Nonresidential zoning	5.1***	4.8***	4.0***	3.5***
	(1.2)	(1.2)	(1.3)	(1.1)
Minimum lot size, median sq ft, ln			−1.1**	0.4
			(0.5)	(0.5)
Distance to downtown, miles	−1.2***	−0.8***	−1.1***	−0.7***
	(0.1)	(0.2)	(0.1)	(0.2)
Minneapolis nearer	−5.2***		2.0	
	(1.7)		(2.4)	
Distance * Minneapolis nearer	0.6***		0.2	
	(0.1)		(0.2)	
Central city	9.5***		11.6***	
	(1.3)		(1.7)	
House size, median sq ft, ln			−5.8***	
			(1.3)	
House value, median $, ln				−13.2***
				(0.8)
Gross rent, median $, ln				−6.7***
				(1.2)
Municipality indicators	No	Yes	No	Yes
Housing age (4 polynomials)	Yes	Yes	Yes	Yes
Observations	2,353	2,345	1,196	2,154
Pseudo R-squared	0.31	0.42	0.44	0.59

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing the non-White population share on zoning and control variables. We inflate the dependent variable by a factor of 100 for ease of reading. Block groups are the unit of observation. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. The natural log of the median minimum lot size is taken over all single-family parcels. Geographic control variables are based on straight-line distance from the block group centroid to the downtowns of Minneapolis and St. Paul. *Minneapolis nearer* is a binary variable that equals 1 for block groups closer to Minneapolis. We interact that binary variable with the distance in miles. *House size* is interior floor space. *Home value* is the tax appraisal of the parcel, including structures and land. *Gross rent* is the monthly rent including utilities. We include median housing age up to a fourth power. Standard errors are reported in parentheses. We allow error correlation to follow an inverse-distance matrix truncated at 4 miles. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.2: Race-Specific Generalized Two-Stage Least Squares Results
Dependent variables: Minority population shares × 100

Regression	Black (5)	Asian (6)	Hispanic (7)	Other (8)
Multifamily zoning	17.6***	−1.5**	4.5***	3.6***
	(1.0)	(0.7)	(0.6)	(0.7)
Middle zoning	5.5***	−0.1	7.4***	7.9***
	(0.9)	(0.6)	(0.5)	(0.6)
Planned Unit Development zoning	6.1***	1.6**	0.3	−0.3
	(1.0)	(0.7)	(0.6)	(0.7)
Nonresidential zoning	2.1**	1.5**	1.1*	1.4**
	(0.9)	(0.6)	(0.5)	(0.6)
Distance to downtown, miles	−0.4***	−0.3***	−0.2***	−0.2***
	(0.1)	(0.1)	(0.1)	(0.1)
Minneapolis nearer	0.0	0.5	0.2	1.6***
	(1.2)	(1.0)	(0.6)	(0.5)
Distance * Minneapolis nearer	0.1	0.0	0.1	0.0
	(0.1)	(0.1)	(0.1)	(0.1)
Central city	6.9***	1.1*	−0.2	0.8
	(0.9)	(0.6)	(0.6)	(0.6)
Housing age (4 polynomials)	Yes	Yes	Yes	Yes
Observations	2,353	2,353	2,353	2,353
Pseudo R-squared	0.28	0.08	0.17	0.20
Group's population share	11%	8%	7%	6%

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing minority population shares on zoning and control variables using generalized two-stage least squares. Rows cannot be summed because errors are presumably correlated across regressions. The bottom row records the seven-county metro population share of each group. If zoning affected each group equally, the coefficients would be proportionate to the population shares. We inflate the dependent variables by a factor of 100 for ease of reading. Block groups are the unit of observation. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. Geographic control variables are based on straight-line distance from the block group centroid to the downtowns of Minneapolis and St. Paul. *Minneapolis nearer* is a binary variable that equals 1 for block groups closer to Minneapolis. We interact that binary variable with the distance in miles. We include median housing age up to a fourth power. Standard errors are reported in parentheses. We allow error correlation to follow an inverse-distance matrix truncated at 4 miles. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.3: Ordinary Least Squares Results
Dependent variable: Non-White population share × 100

Regression	(9)	(10)	(11)	(12)
Multifamily zoning	16.8***	13.3***	20.4***	19.7***
	(4.6)	(5.1)	(4.8)	(3.2)
Middle zoning	18.0***	18.3***	12.3***	9.9***
	(3.0)	(1.8)	(3.0)	(1.3)
Planned Unit Development zoning	9.5***	8.0***	11.7***	6.1**
	(2.8)	(2.5)	(3.7)	(3.0)
Nonresidential zoning	6.7***	5.5***	3.4*	1.3
	(1.9)	(1.4)	(2.0)	(2.6)
Minimum lot size, median sq ft, ln			−2.3***	1.3
			(0.6)	(0.9)
Distance to downtown, miles	−1.6***	−1.5***	−1.4***	−1.5***
	(0.3)	(0.5)	(0.3)	(0.4)
Minneapolis nearer	−7.7**		−5.7	
	(0.0)		(6.7)	
Distance * Minneapolis nearer	0.9***		0.7	
	(0.3)		(0.5)	
Central city	9.1*		18.2***	
	(5.3)		(2.4)	
House size, median sq ft, ln			−8.2	
			(5.0)	
House value, median $, ln				−24.6***
				(6.2)
Gross rent, median $, ln				−7.0***
				(1.5)
Municipality indicators	No	Yes	No	Yes
Housing age (4 polynomials)	Yes	Yes	Yes	Yes
Observations	2,345	2,345	1,196	2,154
R-squared	0.33	0.50	0.47	0.66

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing the non-White population share on zoning and control variables. We inflate the dependent variable by a factor of 100 for ease of reading. Block groups are the unit of observation. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. The natural log of the median minimum lot size is taken over all single-family parcels. Geographic control variables are based on straight-line distance from the block group centroid to the downtowns of Minneapolis and St. Paul. *Minneapolis nearer* is a binary variable that equals 1 for block groups closer to Minneapolis. We interact that binary variable with the distance in miles. *House size* is interior floor space. *House value* is the tax appraisal of the parcel, including structures and land. *Gross rent* is the monthly rent including utilities. We include median housing age up to a fourth power. Standard errors are clustered by municipality and reported in parentheses. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.4: Robustness Results
Dependent variable: Non-White population share × 100

Regression	New block groups only (13)	Suburbs only (14)	Cities only (15)
Multifamily zoning	19.8***	25.3***	17.4***
	(2.8)	(1.6)	(2.9)
Middle zoning	13.0***	12.6***	10.9***
	(2.9)	(1.4)	(2.6)
Planned Unit Development zoning	6.5*	6.2***	
	(1.7)	(1.2)	
Nonresidential zoning	5.7***	4.4***	3.8
	(1.7)	(1.1)	(3.6)
Distance to downtown, miles	−1.2***	−1.2***	−1.9
	(0.2)	(0.1)	(1.5)
Minneapolis nearer	−7.3*	−1.0	−4.6
	(4.0)	(2.6)	(4.8)
Distance * Minneapolis nearer	0.5**	0.4**	−1.1
	(0.2)	(0.2)	(1.5)
Central city	−9.6**		
	(4.1)		
Municipality indicators	No	No	No
Housing age (4 polynomials)	Yes	Yes	Yes
Observations	691	1,698	655
Pseudo R-squared	0.27	0.32	0.19

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing non-White population share on zoning and control variables. In Regression (13), we use only block groups where the median dwelling age is 35 years or less. In Regressions (14) and (15), we use only block groups in the suburbs or central cities, respectively. We inflate the dependent variable by a factor of 100 for ease of reading. Block groups are the unit of observation. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. The central cities have almost no Planned Unit Development zoning and little variation in minimum lot sizes, so we remove those variables for Regression (15). Geographic control variables are based on straight-line distance from the block group centroid to the downtowns of Minneapolis and St. Paul. *Minneapolis nearer* is a binary variable that equals 1 for block groups closer to Minneapolis. We interact that binary variable with the distance in miles. We include median housing age up to a fourth power. Standard errors are reported in parentheses. We allow error correlation to follow an inverse-distance matrix truncated at 4 miles. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.5: Minimum Lot Size Generalized Two-Stage Least Squares Results Dependent variable: Non-White population share × 100

Regression	(16)	(17)	(18)	(19)
Minimum lot size, natural log	−1.1		−5.3***	
	(0.8)		(1.8)	
Maximum dwellings per acre		0.9**		−0.2
(i.e., 43,560/minimum lot size)		(0.4)		(0.4)
Nonresidential zoning	4.4*	4.4*	3.8	
	(2.6)	(2.6)	(2.7)	
Median home value, ln $				−15.9***
				(1.7)
Distance to downtown, miles	−1.1***	−1.0***	−1.1***	−1.3***
	(0.2)	(0.2)	(0.2)	(0.2)
Minneapolis nearer	−0.9	0.6	1.7	1.9
	(3.5)	(3.5)	(3.7)	(3.2)
Distance * Minneapolis nearer	0.2	0.2	0.1	0.1
	(0.3)	(0.2)	(0.3)	(0.2)
Central city	8.4***	7.3***	7.3***	6.6***
	(2.4)	(2.5)	(2.5)	(2.3)
Housing age (4 polynomials)	Yes	Yes	Yes	Yes
Includes large-lot block groups?	Yes	Yes	No	Yes
Observations	479	479	453	479
Pseudo R-squared	0.23	0.23	0.20	0.46

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing non-White population share on median minimum lot size. The sample is limited to block groups in which over 97% of residential land is zoned single-family. In Regressions (16) and (18), we express lot size in log terms. In Regressions (17) and (19), we express it as dwelling units per acre. We inflate the dependent variable by a factor of 100 for ease of reading. Block groups are the unit of observation. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. The central cities have almost no Planned Unit Development zoning and little variation in minimum lot sizes, so we remove those variables for Regression (15). Geographic control variables are based on straight-line distance from the block group centroid to the downtowns of Minneapolis and St. Paul. *Minneapolis nearer* is a binary variable that equals 1 for block groups closer to Minneapolis. We interact that binary variable with the distance in miles. Home value is the tax appraisal of the parcel, including structures and land. We include median housing age up to a fourth power. Standard errors are reported in parentheses. We allow error correlation to follow an inverse-distance matrix truncated at 4 miles. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.6: Non-White Population, Covenants, and Zoning in Hennepin County
Dependent variable: Non-White population share × 100

	Minneapolis only				
Regression	(20)	(21)	(22)	(23)	(24)
ArcSinH (covenant ratio)	−13.2**	−10.7*	−6.7*	−3.6	
	(6.1)	(6.2)	(3.6)	(3.4)	
Covenant indicator		−3.2	−5.5***	−3.7***	
		(2.1)	(1.1)	(1.1)	
Multifamily zoning				25.2***	27.9***
				(2.5)	(2.5)
Middle zoning				14.0***	16.7***
				(2.3)	(2.3)
Planned Unit Development zoning				6.6**	7.8***
				(2.6)	(2.6)
Nonresidential zoning				4.9**	5.2**
				(2.2)	(2.2)
Distance to downtown, miles	−19.8***	−19.5***	−5.1***	−3.1***	−2.9***
	(4.2)	(4.2)	(0.8)	(0.9)	(1.0)
Distance to downtown, squared	2.0***	2.1***	0.1***	0.1***	0.1**
	(0.6)	(0.6)	(0.0)	(0.0)	(0.0)
Municipality indicators	No	No	Yes	Yes	Yes
Housing age (4 polynomials)	Yes	Yes	Yes	Yes	Yes
Observations	387	387	1,048	1,048	1,048
Pseudo *R*-squared	0.10	0.10	0.34	0.34	0.32

Sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing non-White population share on racial covenant and zoning data. We use generalized two-stage least squares and allow errors to be correlated following an inverse-distance matrix truncated at 4 miles. The sample is limited to the city of Minneapolis in Regressions (20) and (21) and Hennepin County in the rest. We inflate the dependent variables by a factor of 100 for ease of reading. Block groups are the unit of observation. We include covenant data in two ways: as the hyperbolic inverse sine of the number of recorded covenants divided by the number of current residential tax parcels and as a binary variable that equals 1 in block groups with at least 1 recorded covenant. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. Geographic control variables are based on straight-line distance from the block group centroid to downtown Minneapolis. We include median housing age up to a fourth power. Standard errors are reported in parentheses. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.7: Black Population, Covenants, and Zoning in Hennepin County
Dependent variable: Black population share × 100

	Minneapolis only				
Regression	(25)	(26)	(27)	(28)	(29)
ArcSinH (covenant ratio)	−8.0	−7.2	−2.9	−0.9	
	(5.1)	(5.2)	(2.7)	(2.6)	
Covenant indicator		−1.0	−3.3***	−1.8**	
		(1.8)	(0.9)	(0.9)	
Multifamily zoning				21.0***	21.6***
				(1.9)	(1.9)
Middle zoning				7.6***	8.3***
				(1.7)	(1.7)
Planned Unit Development zoning				6.0**	6.4***
				(1.9)	(2.0)
Nonresidential zoning				2.0	2.2
				(1.7)	(1.7)
Distance to downtown, miles	−11.6***	−11.7***	−4.5***	−2.1***	−1.9**
	(3.6)	(3.6)	(0.7)	(0.8)	(0.8)
Distance to downtown, squared	0.9*	1.0*	0.2***	0.1**	0.1**
	(0.5)	(0.5)	(0.0)	(0.0)	(0.0)
Municipality indicators	No	No	Yes	Yes	Yes
Housing age (4 polynomials)	Yes	Yes	Yes	Yes	Yes
Observations	387	387	1,048	1,048	1,048
Pseudo R-squared	0.14	0.14	0.27	0.28	0.28

Data sources: Various including 2020 Census; jurisdiction zoning and tax data; American Community Survey (ACS) 2019 5-year aggregates.

Note: This table reports the results of regressing black population share on racial covenant and zoning data. We use generalized two-stage least squares and allow errors to be correlated following an inverse-distance matrix truncated at 4 miles. The sample is limited to the city of Minneapolis in Regressions (25) and (26) and Hennepin County in the rest. We inflate the dependent variables by a factor of 100 for ease of reading. Block groups are the unit of observation. We include covenant data in two ways: as the hyperbolic inverse sine of the number of recorded covenants divided by the number of current residential tax parcels and as a binary variable that equals 1 in block groups with at least 1 recorded covenant. The three residential zoning categories are expressed as shares of residential land. Nonresidential zoning is a share of all zoned land. Geographic control variables are based on straight-line distance from the block group centroid to downtown Minneapolis. We include median housing age up to a fourth power. Standard errors are reported in parentheses. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively.

Appendix Table A.8: Observables Used in Blinder–Oaxaca Decomposition

Household income, ln	Disability status (binary)
Educational attainment (5 categories)	Veteran status
Child presence	Employment status
Immigrant presence	Employment status * age > 64
Head noncitizen status	Sex
Marital status	Age and square root of age

Appendix Table A.9: Ownership Gaps across Major Metro Areas

	Ownership rate (%)			Unexplained gap (percentage points)			
Metro area	White	Black	Asian	Hispanic	White–Black	White–Asian	White–Hispanic
Minneapolis	76	25	59	45	26	4	9
Pittsburgh	74	33	47	43	22	11	11
Chicago	75	41	64	53	22	4	6
Detroit	78	43	62	60	21	11	3
Phoenix	71	34	64	51	21	3	3
St. Louis	77	40	57	56	21	8	4
Cleveland	75	36	54	44	20	13	10
Cincinnati	72	34	56	44	20	11	5
Charlotte	75	44	63	45	19	6	8
New York	67	33	53	28	19	−2	16
Tampa	72	40	66	51	19	3	8
Baltimore	77	46	66	51	18	3	6
Riverside	72	42	72	57	18	−2	0
Nashville	73	43	58	41	18	5	4
Denver	70	40	62	48	18	0	5
Indianapolis	73	38	57	46	17	8	3
Las Vegas	63	30	63	45	17	−1	0
Seattle	65	31	60	39	17	−3	4
Sacramento	67	34	63	46	17	−5	3
Atlanta	76	47	68	47	17	1	5
Kansas City	71	38	57	51	16	1	−1
Dallas	69	37	63	52	16	0	−3
Columbus	68	34	51	40	16	4	3
Houston	72	42	69	53	16	−3	−2
San Diego	61	30	58	39	15	−3	4
Miami	73	46	70	51	15	−5	10
Boston	69	36	54	28	14	−1	12
Portland	66	35	66	40	13	−3	5
Orlando	71	45	74	48	13	−7	10
San Antonio	70	45	55	58	13	4	1
Washington, DC	72	51	69	49	13	−4	2
San Francisco	61	34	60	39	13	−7	5
Philadelphia	76	49	64	45	12	1	7
Los Angeles	57	34	56	38	11	−6	−1
Austin	65	42	57	48	11	−2	1

Sources: American Community Survey (ACS) 2019 5-year data; Ruggles et al. (2021).

Note: This table displays homeownership rates and the unexplained portion of the gaps between minority and White households. Households are categorized by the race or ethnicity of household heads and the data cover the 35 largest U.S. metropolitan areas. We used a Blinder–Oaxaca decomposition to identify the portion of the minority–White gap that can be explained by the observable characteristics listed in Appendix Table A.8. Our decomposition is based on the rate. We lack data on credit characteristics and parental wealth, which would explain still more of the gap. The last three columns show the ownership gap that cannot be explained by differences in observable characteristics. In some cases, observable data overexplain the gap, leaving a negative unexplained portion.

Appendix Table A.10: Single versus Multifamily Occupancy
Binary dependent variable: Household occupies detached house

	United States		Twin Cities (7-county metro)	
	Owners	Renters	Owners	Renters
Regression	(30)	(31)	(32)	(33)
Black	0.44***	0.50***	−0.09	−0.28
	(0.08)	(0.04)	(0.38)	(0.21)
Asian	−0.21**	−0.28***	−0.77	−0.06
	(0.10)	(0.06)	(0.55)	(0.35)
Hispanic	0.07*	−0.19***	−0.30	0.32
	(0.04)	(0.03)	(0.33)	(0.32)
Other	0.10	−0.19***	0.36	−0.09
	(0.07)	(0.03)	(0.33)	(0.27)
Immigrant present	−0.30***	−0.16***	−0.27	−0.33
	(0.03)	(0.03)	(0.20)	(0.23)
Black*Immigrant	0.01	−0.41***	−0.25	−0.71**
	(0.10)	(0.06)	(0.73)	(0.33)
Asian*Immigrant	0.28***	−0.20***	0.53	−0.24
	(0.10)	(0.07)	(0.72)	(0.33)
Hispanic*Immigrant	0.05	−0.01	0.06	−0.44
	(0.06)	(0.04)	(0.78)	(0.43)
Other*Immigrant	−0.07	−0.09	—	0.87
	(0.10)	(0.07)	—	(0.76)
Other controls	Yes	Yes	Yes	Yes
PUMA fixed effects	Yes	Yes	Yes	Yes
Observations	12,976,118	11,522,456	326,968	224,365
Pseudo *R*-squared	0.20	0.13	0.20	0.19

Note: This table reports the result of a logistic regression on housing typology. The dependent variable equals 1 for a household occupying a single-family detached house and 0 for a household occupying a dwelling in a building with three or more units. Households in other typologies, including most middle housing occupants, are excluded. Positive coefficients imply higher odds of occupying a detached house. The sample includes households in public use microdata areas (PUMAs) in which both detached and multifamily housing account for at least 30% of units. The first two regressions cover national data; the latter two cover the Twin Cities. In each geography, we further segment the sample and consider owners and renters separately. Standard errors are reported in parentheses. Race variables refer to the household head. *Immigrant* is a binary variable that equals 1 for households in which any member was born outside the United States. Statistical significance is reported at the 10%, 5%, and 1% levels with *, **, and ***, respectively. Other controls include income, child presence, marital status, sex, educational attainment, disability status, veteran status, age, employed, and an interaction term between employed and age above 65.