Public Housing Preferences: Evidence from New York City 1930-2010*

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Abstract

This paper studies the long-run effects of public housing construction on neighborhood composition and housing markets in New York City from 1930 to 2010. Using newly assembled archival data linked to census tract information, I track racial population changes separately for public and private residents, allowing me to isolate behavioral sorting from mechanical placement. Employing a difference-in-differences shows that projects built before 1970 triggered persistent market responses: private White and Black populations declined by 88% and 76%, while rents fell by up to 43%, consistent with declining neighborhood demand. In contrast, post-1970 projects—built under different policy regimes—had muted effects, driven largely by the direct placement of public housing tenants. To quantify preferences over public housing design, I use a discrete choice model of location to map reduced-form coefficients into marginal willingness to pay estimates. The results show that taller buildings depress demand, higher construction quality increases it, and resident composition shapes preferences in racially patterned ways. These results highlight a trade-off for policymakers: expanding affordable housing in high-opportunity areas requires balancing scale with quality to mitigate adverse market responses.

JEL codes: I31, R21, R28.

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1 Introduction

Throughout the 20th century, Public housing has long been a cornerstone of housing policy in the United States, providing shelter to low-income households and reshaping urban neighborhoods. Yet, beyond its role as a safety net, public housing also acts as a durable, place-based intervention that alters the trajectory of neighborhoods through its design, location, and the profile of incoming residents. At its onset in the mid-1930s, federal public housing in the United States was designed more than just a safety net—it was a highly visible, place-based policy that reflected and reinforced social priorities through its design, tenant selection, and spatial footprint (Goetz, 2012; Bloom et al., 2016).These developments were often greeted with broad public support and celebrated as improvements over slum conditions. But by the late 1960s, cracks had begun to appear in this narrative.

First, the physical form of public housing—most notably the *Tower in the Park* design criticized by Jane Jacobs and Oscar Newman—was increasingly seen as unsustainable, fostering isolation, crime, and disinvestment (Jacobs, 1992; Newman, 1997; Plunz, 2016). Second, public housing was never race-neutral. Early projects often reinforced patterns of racial segregation, both in their physical siting and in tenant screening procedures that prioritized families with stable incomes and employment.[1] In cities like New York, this translated into racially designated developments, limiting access for the poorest residents and reflecting broader social exclusion (Allen and Van Riper, 2020). Yet, despite its scale and permanence, the literature takes public housing a homogenous policy and we know surprisingly little about how heterogeneous characteristics of public housing affects surrounding residents — and how these effects vary depending on what gets built and for whom.

This paper investigates how the construction of public housing reshapes neighborhood composition and housing markets over time. I provide new evidence on how different ethnic groups respond to public housing developments in their vicinity, and specific design characteristics — such as building height, construction quality, and tenant composition — are being valued as by private market residents.

To examine the long-run effects of public housing on neighborhood composition and housing markets, I compile a new dataset linking archival records on all 275 public housing projects in New York City to decennial U.S. Census data from 1930 to 2010. I track tract-level population counts by race and ethnicity—White, Black, and Hispanic—and merge them with detailed information on project location, construction date, and tenant composition at opening and over time. This allows me to distinguish sorting by private residents

[1] Public housing authorities prioritized households with a 'nuclear family' structure — married, male-headed, and with children — while racial segregation and selective tenant policies shaped demographic composition. Citizenship and nativity also influenced whom officials deemed deserving of public benefits. For example, between 1935 and 1944, the Public Works Administration allocated about one-third of its public housing projects to only African Americans (Radford, 2008).

from mechanical effects driven by public housing placement. I further collect rich project-level attributes—construction costs (a proxy for quality), building height, ground coverage, and initial racial composition—to study how design and demographics shape neighborhood responses.

To estimate effects on housing markets, I use tract-level median contract rents from the U.S. Census. To validate and complement these measures, I assemble a novel dataset of 19,446 property-level asking rents from the New York Times real estate section, digitized at ten-year intervals from 1930 to 2010. These data allow for a more granular assessment of market dynamics and help address concerns about compositional bias in census-reported rents.

I identify the causal effect of public housing by leveraging the staggered implementation of public housing projects across the city. Importantly, the areas targeted by the NYC Housing Authority to build public housing are not random. Thus, I utilize the distance to public housing projects as a measure of treatment intensity, allowing for the estimation of disparate effects on market rents, Black and White populations. Specifically, I assign the treatment at the census tract level and compare tracts with a public housing project to slightly more distant tracts (Blanco and Neri, 2025).

My findings show that public housing construction led to large, persistent changes in neighborhood composition. In the decade following construction, the total White population in treated tracts initially rose by 46.2%, but by 60 years later had declined by 44.1% relative to baseline. In contrast, the Black and Hispanic populations increased substantially and remained elevated, reaching 160% and 175.1% of baseline levels, respectively, six decades post-construction.

Crucially, by separating public housing residents from the rest of the population, I am able to identify market-driven responses among private households. I find that the number of privately housed White and Black residents declined by 49.2% and 36.9% within 10 years, and by 87.9% and 75.9% over the long run—corresponding to net losses of approximately 2,997 White and 1,141 Black residents per tract. In contrast, there is no significant effect on the net Hispanic population, suggesting that most of the observed increase was due to direct placement into public housing rather than sorting in the private market. These patterns indicate that the observed White outflows reflect residential flight, while the sustained growth of Black and Hispanic residents was driven primarily by public housing allocations.

These population shifts were accompanied by substantial declines in local rents. Median contract rents fell by 42.7% over 60 years. I find similarly sharp declines in market rents, using rental listings from the New York Times, : median and average rent per room declined by 67.8% and 75.0%, respectively. Taken together, the concurrent decline in rents and the out-migration of privately housed residents provide evidence that public housing construction reshaped the desirability of the surrounding neighborhoods in the private market.

A central insight from the analysis is the pronounced heterogeneity in the effects of public housing across construction periods. Projects built before 1970 generated large and persistent changes in neighborhood composition and housing markets. These 171 early developments triggered substantial outflows of privately housed White and Black residents of 489 and 1,026, sizable rent declines of 12.28% 30 years post construction. These findings reinforce the conclusion that early public housing construction had a lasting impact, primarily through private market responses.

In contrast, post-1970 projects had far more limited effects. Population changes reflect the administrative placement of residents into public housing units, while the effect on neighborhood remains less well estimates or clear. This stark divergence underscores that public housing is not a uniform treatment. The comparison highlights a fundamental shift in the design and consequences of public housing policy after 1970, from an urban development tool that reshaped entire neighborhoods to a narrowly targeted program.

I use a static equilibrium model of neighborhood choice, following Bayer, Ferreira, et al. (2007) and Almagro et al. (2023), to translate reduced-form estimates into individual preferences for living near public housing. By mapping the model's comparative statics to the empirical coefficients, I recover preference parameters and their associated variance, leveraging the sampling variance of the underlying reduced-form estimates.

The design and construction quality of public housing significantly influence neighborhood desirability over time.[2] Survey evidence supports this view: the 2003 "Housing Illinois" survey found that 47% of residents perceived low-income housing designs as unattractive (Belden and Russonello, 2003). Using the structure of the model and empirically recovered preference parameters, I estimate how much additional rent households are willing to pay in response to differences in public housing attributes.

I show that public housing design plays a key role in shaping neighborhood preferences. Higher construction costs — used as a proxy for build quality — are consistently associated with higher marginal willingness to pay (MWTP) among all racial and ethnic groups living in the private housing market. For a $1,000 increase in construction costs, White, Black, and Hispanic households are willing to pay up to $40–$54 more per month to live in nearby neighborhoods in the long run.

By contrast, larger-scale developments are perceived as less desirable. A 100-unit increase in the number of public housing apartments generates sizable negative MWTPs across groups, particularly among Black and White households, who would require between $200–$280 per month in compensation to remain in affected tracts. These results suggest

[2] Architecture as an amenity remains a sparsely studied topic. Ahlfeldt and Holman (2018) find that properties in architecturally distinctive areas command a price premium.

concerns about density, stigma, or the depressive effects on nearby rents.

Design features also matter. Taller public housing buildings—central to the Tower in the Park model—are consistently viewed as undesirable. A one-story increase in height reduces long-run MWTP by \$122.74 for White residents and \$126.93 for Black residents, with additional spillover effects in adjacent tracts, particularly among Black households. These findings align with longstanding critiques by urban thinkers like Jane Jacobs, who argued that high-rise structures surrounded by open space foster unsafe and socially isolating environments (Jacobs, 1992). In contrast, ground coverage — a proxy for compact, low-rise design—has no consistent effect on MWTP across racial groups, suggesting that vertical scale, rather than spatial openness per se, is the primary design element shaping neighborhood desirability.

Finally, I document sharp racial asymmetries in preferences for the racial composition of public housing tenants. White residents strongly prefer proximity to public housing developments with growing shares of White tenants: a 10 percentage point increase in the White tenant share from project opening to 2010 raises their MWTP by \$254.78 in the long run. In adjacent areas, the increase is \$164.55. Black residents, by contrast, avoid proximity to developments with increasing shares of Black tenants, requiring a monthly discount of \$369.63 in the long run. Instead, they express a positive MWTP of \$213.26 for projects with growing Hispanic tenant shares.

Hispanic residents show weaker responses overall. Their MWTP estimates for racial composition are generally imprecise and statistically insignificant, but one robust pattern emerges: in the long run, a 10 percentage point increase in the Black tenant share reduces their MWTP by \$116.41. This suggests some aversion to living near public housing developments with growing Black populations, but otherwise little evidence of strong racial sorting behavior among Hispanic households.

These findings demonstrate that private households prefer well-built, smaller-scale public housing, but respond sharply to both architectural design and the racial composition of residents. This highlights a core policy tradeoff: expanding affordable housing in high-opportunity areas requires balancing scale and acceptance. Building more units may provoke adverse market reactions, while smaller, higher-quality developments are more broadly accepted but constrain the overall scale of affordable housing provision.

A central contribution of this paper is the analysis of heterogeneity in public housing attributes, including project height, construction quality and other public housing characteristics. While prior research has examined the average effects of affordable and public housing on neighborhood outcomes, I focus instead on how the physical characteristics of these developments shape individual preferences about public housing design. I focus on the largest urban housing market in the United States, New York City, although the effects should generalize to other cities with large historical public housing policies, such as Chicago.

Unlike most of the literature, which studies the effects of public housing demolition, this paper provides the first causal, quantitative analysis of public housing construction over an 80-year period.[3] By comparing construction cohorts, I show that the neighborhood effects of public housing depend critically on the design and quality of the developments themselves. These findings offer a new perspective on why earlier, high-rise projects often had negative long-run impacts, while better-designed, smaller-scale developments were more positively received.

In the context of public housing demolitions in Chicago, previous studies have identified significant positive effects. The demolition of public housing has been associated with substantial increases in nearby home prices (Almagro et al., 2023; Blanco, 2022), significant reductions in crime (Aliprantis and Hartley, 2015; Sandler, 2017), and notable shifts in neighborhood socioeconomic composition (Tach and Emory, 2017; Blanco, 2022). In contrast, in New York City, subsidized housing has generated significant price appreciation in the immediate vicinity (Schwartz et al., 2006). Federal public housing constructed between 1977 and 2000 has not typically led to reductions in property values (Ellen et al., 2007).

Moreover, the price effects of affordable housing construction depend on neighborhood composition. Diamond and McQuade (2019) show that Low-Income Housing Tax Credit developments lead to price increases in low-income neighborhoods but cause price declines in high-income areas.[4] Revitalization projects or the transition to mixed-income housing can attract higher-income residents in low-income areas (Blanco and Neri, 2025; Staiger et al., 2024). Moreover, aside from Diamond and McQuade (2019) and Almagro et al. (2023), few studies apply a theoretical framework to put structure on individual preferences for public housing.

More broadly, my paper is related to the literature that examines the spillovers to neighborhoods of housing policies. Rossi-Hansberg et al. (2010) analyze the impact of an urban revitalization program in Richmond, Virginia, finding that targeted investments led to 2–5% annual land price increases, with housing externalities diminishing by half every 1,000

[3]The study of public housing demolition in Chicago dates back to the early stages of initiatives like the Moving to Opportunity projects. However, this body of literature estimates the demolition effect on individual-level outcomes rather than neighborhoods. Studies find moderately positive impacts from public housing demolition, particularly for residents, and minority populations (Jacob, 2004; Chetty et al., 2016; Chyn, 2018). More recent research has extended its scope to examine the consequences of demolitions on a broader range of outcomes, including rental rates and construction trends.

[4]Additionally, beyond the direct neighborhood effects of social housing, a broader literature examines the labor market consequences of these public housing availability. For example, Dalmazzo et al. (2022) and Bromhead and Lyons (2022) study the effects of historical housing policies on population dynamics, labor supply, and industry location. Baum-Snow and Marion (2009) and Eriksen and Rosenthal (2010) find significant crowd-out effects of Low-Income-Housing-Tax-Credit developments on new market-rate housing supply.

feet. Redding and Sturm (2024) examine the long-term effects of World War II bombing in London as an exogenous shock to neighborhood composition and property prices. They show that poor-quality post-war construction led to a local decline in post-war property values and a decrease in the share of high-income residents with significant spillover effects on surrounding areas. Asquith et al. (2023) examine the effects of large new apartment buildings on nearby housing prices and neighborhood composition. Similarly, Campbell et al. (2011) examine the effects of housing foreclosure. Autor et al. (2014, 2017) study the impact of ending rent control on nearby real estate prices and crime rates. I provide new evidence on neighborhood effects from public housing. Closely related to the findings by Redding and Sturm (2024), I show that higher-quality buildings had positive long-run effects on the MWTP for these neighborhoods.

Finally, a large body of research shows that households pay premiums for neighborhood amenities, including demographic composition and social context (Bayer, Ferreira, et al., 2007; Bayer, McMillan, et al., 2016; Diamond and McQuade, 2019). In the case of public housing, which is shaped by racialized allocation patterns, such preferences can reinforce residential segregation (Schelling, 1971; Card et al., 2008; Logan and Parman, 2017). In this paper, I incorporate these dynamics by explicitly modeling preferences over neighborhood racial composition. Building on Bayer, Ferreira, et al. (2007) and Almagro et al. (2023), I use reduced-form estimates from NYC's 80-year public housing rollout to recover group-specific preference parameters. The results reveal strong evidence of racial homophily: White households have positive MWTP for areas with increasing shares of White public housing residents, while Black residents are negatively valued by nearly all groups. By contrast, households express weak or no systematic preferences regarding Hispanic residents.

The paper proceeds as follows. Section 2 outlines the historical context. Section 3 describes the data sources, and Section 4 presents the empirical strategy. In Section 5, I estimate the long-run effects of public housing, and Section 6 examines heterogeneity across project construction cohorts. Section 7 introduces the theoretical model and describes the estimation of preference parameters. Section 8 reports marginal willingness to pay estimates, and Section 9 concludes.

2 Background

Public housing emerged as a key response to urban poverty during the mid-20th century, aiming to replace slums, stabilize communities, and provide affordable housing to working-class families. These efforts also reshaped neighborhood racial and socioeconomic dynamics in lasting ways. In response to the Great Depression's economic devastation and housing crisis, federal programs like the Public Works Administration and the 1937 Wagner-Steagall

Act launched a wave of slum clearance and new construction to foster health and safety through modern housing (Allen and Van Riper, 2020; Radford, 2008; Fogelson, 2003).

In New York City, the newly formed Housing Authority (NYCHA) pioneered high-quality, professionally managed public housing. Its first project, the First Houses (1935), exemplified the era's optimism: replacing overcrowded tenements with clean, modern apartments for working families (Hunter College, 2025). Early residents often recalled public housing as "paradise," illustrating the broader optimism and strong community cohesion that marked the early era of federally supported housing. (Williams, 2014; Bloom, 2008; Marcuse, 1986; Hunt, 2009).[5]

Authorities like NYCHA and the Chicago Housing Authority implemented strict tenant screening, emphasizing stable employment, nuclear families, and proximity to transit and schools. Early projects often reinforced patterns of racial segregation, with tenant selection policies and housing placements reflecting broader societal prejudices. In New York City, some developments were designated for White families, while others primarily housed Black or Hispanic households.[6] These developments were operated like private apartment complexes, with routine maintenance and community-building efforts (Hunt, 2009; Bloom, 2008).

Demand for public housing surged in the postwar period amid a housing shortage and returning servicemen. The 1949 Housing Act spurred expansion while reaffirming its dual goals of slum clearance and affordability. In NYC alone, 72,499 units were built in the 1950s and 42,721 more in the 1960s (Plunz, 2016). Construction sites were typically selected based on land costs and surrounding population densities, leading to a concentration of developments in lower-cost neighborhoods.[7] By 1970, public housing made up 25% of all new residential construction. Designs shifted from low-rise blocks to high-density towers in the modernist *Tower in the Park* model, emphasizing open space around high-rise buildings.[8]

[5]Maude Davis, a retired public school principal, described Altgeld Gardens in Chicago - opened in 1945 - as "paradise," noting, "We felt this was just the greatest housing that we could live in! There was pride in it." So did Vonsell Ashford when moving into the 1955 completed Harold L. Ickes Homes in Chicago: "The building was new, and they had a beautiful playground for the children. You couldn't ask for a better location, and the place was just marvelous. I had three bedrooms, a nice storage area, and a linen closet ... I thought I was moving to paradise" (Hunt, 2009). Though expressed by Chicago residents, these sentiments reflect the broader optimism that characterized early public housing, including NYCHA developments.

[6]Moreover, early public housing projects exhibited explicit segregation, with a predominant allocation to White residents (see Figure C2.3). Against this backdrop, tracts with public housing projects followed different demographic trajectories than trends in the rest of New York City. Specifically, tracts designated for public housing had significantly higher White population in 1930. In 2010, the same tracts had a lower White and higher Black population than the average tract in the rest of the city (see Figure C1.1).

[7]Federal regulations capped land acquisition costs at \$1.50 per square foot, making many centrally located slum areas financially infeasible. NYCHA faced considerable challenges assembling large parcels in built-up areas due to fragmented ownership, political opposition, and zoning constraints. As a result, the Authority often settled on sites that were affordable but marginal, including deindustrialized zones and peripheries of established neighborhoods (Genevro, 1986).

[8]Figure C2.1 shows the evolution of public housing construction by construction cohort.

By 1968, NYCHA eased its admission criteria, declaring that it would no longer evaluate applicants based on moral considerations. Landmark cases such as *Escalera v. NYCHA* (1970) led to the relaxation of strict tenant screening practices, including the elimination of race-based admissions policies that had long characterized early projects. While these changes reflected broader efforts to promote equity, they also contributed to increasing concentrations of poverty within developments. This period coincided with a decline in federal funding and growing maintenance backlogs, straining NYCHA's resources and infrastructure (Bloom, 2008).

The trajectory of public housing in New York City diverged sharply from its early promise. The optimism embodied in the initial construction boom gradually gave way to concerns about crime, neglect, and physical deterioration. The public housing design, called the *Tower in the Park*, drew criticism and lost public support. Famously, urban sociologist Jane Jacobs and architect Oscar Newman criticized the *Tower in the Park* as a utopian design that created places marked by crime and social isolation—spaces that were difficult to monitor and maintain. Rising opposition resulted in a policy shift at the local level in favor of low-density public housing (Jacobs, 1992; Newman, 1997).

Despite these challenges, NYCHA maintained a degree of resilience that distinguished it from other public housing systems, largely due to its commitment to management and maintenance (Bloom, 2008). However, NYCHA's developments were not immune to broader urban challenges. As Marilyn Jones, who moved into the Queensbridge housing project in 1970, puts it: "In the beginning for about the first 2 or 3 years it was fine, but then all of a sudden, crime started, people running around with guns, shooting everybody, people throwing people off the rooftops, police all over the place everywhere." (Petrus and Rosner, 2019).

By the 1980s, chronic underfunding and demographic shifts exacerbated social and economic isolation within public housing communities. Federal programs like HOPE VI aimed to modernize or demolish aging public housing and introduce mixed-income communities (Goetz, 2012; Fernandez, 2010). The 2003 "Housing Illinois" report documents concerns among private residents about public housing's poor maintenance (66%), increased crime (52%), declining property values (49%), and unattractive design (47%). These sentiments echoed in New York City, fueled growing resistance to public housing developments (Belden and Russonello, 2003).

Today, NYCHA remains emblematic of both the early promise and enduring challenges of public housing. Its developments house over 400,000 residents, making it the largest public housing authority in North America. However, it faces a $40 billion capital repair backlog, with aging infrastructure and funding constraints threatening its future. Nationally, public housing continues to grapple with questions of sustainability, equity, and integration

(NYCHA, 2021).

3 Data

This section describes the data used to estimate the effect of public housing on neighborhood composition and preferences. First, I outline the sources and information on public housing, including project locations, construction dates, and key characteristics. Next, I introduce the three key equilibrium variables, as motivated by the theoretical framework in Section 7: market rents, and Black and White population.

Public Housing Characteristics. I obtain information from three sources. First, I utilize the New York City Public Housing Administration (NYCHA) Development Data Book, available annually from 1948 to today. It provides information on construction date, height, number of apartments, construction costs,[9] and ground coverage—the total ground floor area of a project's building footprints divided by the project's total area. Information for the year 1940 in the NYCHA Development Data Book is inferred from archival sources from the Wagner and LaGuardia Archives. In total, there are 299 projects operated by NYCHA.[10] Second, I supplement this dataset with information on the racial composition of public housing projects, including the number of White and Black residents at the time of their initial opening and the racial composition in each census year. For developments constructed up to 1971, I digitize data from the Wagner and LaGuardia Archives. An example of race statistics is shown in Figure C2.4. For more recent years, I use the NYCHA Resident Data Book. Since I evaluate the effect of public housing on neighborhoods, I spatially match public housing projects to 2010 census tracts. Due to their size, some projects span multiple tracts. To account for this, I weight demographics, apartments, and ground coverage proportionally to each project's area share.

Rents. One of the three key equilibrium variables is rent prices. I primarily use the median contract rent reported in the decennial census, which is available from 1930 to 2010 at the census tract level. However, this variable has several limitations. First, it is top-coded, meaning that rents above a certain threshold are reported as a single value, which

[9]Construction costs exclude land acquisition expenses, which are accounted for separately as part of development costs.

[10]Since I am interested in the construction impact, I disregard redevelopment projects that occurred after 1964. Moreover, not all NYCHA projects were built by public entities. Some were developed under the *Turnkey* program, where private developers purchased and constructed buildings, which NYCHA later acquired. While I include these types of projects, this method was introduced in 1969 and accounts for 76 developments.

can distort the true distribution of rents. Second, it might not reflect true market rents, as it includes public housing rents, which are often subsidized.

Therefore, I complement the census data with private rental market data from the *New York Times* real estate section. I digitize rent prices from the *New York Times* real estate section for each decennial census year from 1930 to 2010 to examine the impact of public housing on rents resulting in 19,446 rental listings. Only properties with an exact address or cross-street information were included to ensure accurate geolocation, and the Google Maps API was used to geocode the rental data. Using property-level rent data presents both advantages and limitations. On the one hand, it circumvents issues inherent to median contract rent reported in the United States census; on the other hand, relying on newspaper data, particularly from the *New York Times*, introduces two key limitations. First, as an upper-middle-class newspaper, the *New York Times* may not provide comprehensive coverage across all market segments and is biased toward the higher end of the market. However, no newspaper systematically covers the lower end of the rental market. Table C4.1 compares both rent measures and shows that listing prices are considerably higher than census rents. However, when comparing the indexed median contract rent to both the average real rent per room from NYT listings and a hedonic rent index constructed from listing characteristics, all three measures exhibit similar trend growth over time (see Figure C4.2). Moreover, the average rent per room closely aligns with the level trend of the median contract rent. Since the empirical strategy relies on a difference-in-differences design, it is the evolution of rents over time — rather than differences in levels — that matters for identification. Second, the listing data have limited spatial coverage, as not all tracts are consistently represented in each year.[11] This limitation makes the census data more advantageous for tract-level analysis, given its comprehensive coverage across all tracts and years. A full description of the data collection procedure, summary statistics, and an example of the source material are provided in Appendix C4.

Demographic Information. The last three equilibrium variables are White, Black, and Hispanic populations. I obtain population data as well as census tract boundaries from 1930 to 2010 from IPUMS and IPUMS-NHGIS (Manson et al., 2024). A challenge when building a geographically consistent panel dataset is the presence of boundary changes over time. Census tract boundaries underwent substantial changes throughout most of the 20th century, especially in Brooklyn (Kings County) and Queens County.

Therefore, I harmonize tract boundaries from 1930 to 1950 by adjusting them to 2010 census tract definitions using overlapping area weights, allowing for the construction of a balanced panel. For 1930, I rely on full-count population census data, aggregated at the

[11] Figure B.4 displays the location of the geocoded listings by year.

enumeration district level, which is geographically much smaller than census tracts. Mapping these enumeration districts to 2010 census tracts helps reduce potential measurement error. A potential limitation of this approach, however, is the implicit assumption that tract-level population characteristics are uniformly distributed across space. Further details and data sources are provided in Appendix C1.[12] Finally, I compute White, Black, and Hispanic population net of public housing residents using the information from the LaGuardia and Wagner Archives and the NYCHA Residents Data Book. For each tract I aggregate the total number of public housing residents, which I then subtract from the tract's total population of the respective ethnic group.

Sample The final sample consists of a panel of 2164 census tracts based on 2010 tract boundaries per year from 1930 to 2010. The final set has 225 public housing tracts and ca. 1,500 rental observations per year. All prices and costs have been deflated using the CPI deflator and normalized to the 2010 CPI level. Descriptive statistics for public housing can be found in Appendix C2 and detailed rental statistics are given in Appendix C4. The following section describes the empirical strategy to estimate the causal effects of public housing and further transformations of the data.

4 Empirical Strategy

This section describes the empirical strategy for estimating the causal effects of public housing on population demographics and rent prices. I employ a difference-in-differences (DiD) approach, leveraging variation in the timing of public housing construction and proximity to treated tracts.

Empirically, the approach follows a standard two-way fixed-effects (TWFE) model. Treatment is assigned at the tract level based on whether a tract contained a public housing project at any point within a census year. The completion date serves as the event triggering the treatment effect, as is common in the literature (Asquith et al., 2023; Pennington, 2021).[13]

$$y_{i,t} = \beta \mathbb{1}\left(t \geq Y\right) + \gamma_i + \gamma_t + u_{m,t} \tag{1}$$

Where γ_i denotes group fixed effects and γ_t denotes census year fixed effects, $\mathbb{1}\left(t \geq Y\right)$

[12] I also harmonize the median contract rent and the median household income from the census data to 2010 census tract boundaries. Because income is not a count variable, I calculate tract-level estimates as weighted averages across historical geographies using overlapping area weights.

[13] Since I use decennial census years, projects completed within a given decade first appear as treated at the end of that period. For example, projects completed between 1961 and 1970 are observed as treated in 1970. Consequently, treatment effects in any given census year reflect the weighted average effect of all projects completed within the corresponding treatment year cohort.

is an indicator variable equal to one for treated tracts if census year t is larger than the treatment year Y. Moreover, the coefficient β corresponds to the equilibrium comparative statics for Black and White population, and rent prices, Equation 7 and Equation 6.

The main challenge in the empirical analysis is selecting a suitable comparison group that accurately reflects what would have occurred in the absence of public housing. Ideally, one would conduct an experiment randomly assigning public housing projects to census tracts. However, such an experiment is not feasible. Instead, I must address the concern that the allocation of public housing across the city can be correlated with pre-construction tract and household characteristics. For example, construction sites were chosen based on the price of land and population density, which makes such tracts more likely to be selected for construction than those without.

To address this challenge, I utilize a stacked difference-in-differences design following Blanco and Neri (2025) that uses the variation in proximity to public housing projects to define the comparison group. I create rings of census tracts around each treated tract to define proximity and construct two rings of tracts around each treated tract. The outer ring serves as the comparison group to treated tracts and tracts in the inner ring. Because tracts have fixed boundaries, proximity is defined by being adjacent to a public housing project. Treated tracts have been excluded from any other first or second ring, ensuring that each treated tract's control group consists only of never-treated tracts. Doing so for each project requires appending these tract-project rings such that tracts may occur several times in the dataset. Figure 1 illustrates the spatial layout of fixed tract rings and overlapping tracts. The key assumption is that, in the absence of public housing, demographics and rents would change similarly in both the treated tract and the tracts in the control group. Any differences in outcomes should only be due to the impact of public housing.

The validity of this strategy requires balanced demographics and rents across control and treatment groups prior to treatment. For census outcomes, I estimate the following event study equation at the census tract k, neighborhood (NTA) m, project p, and year t level:

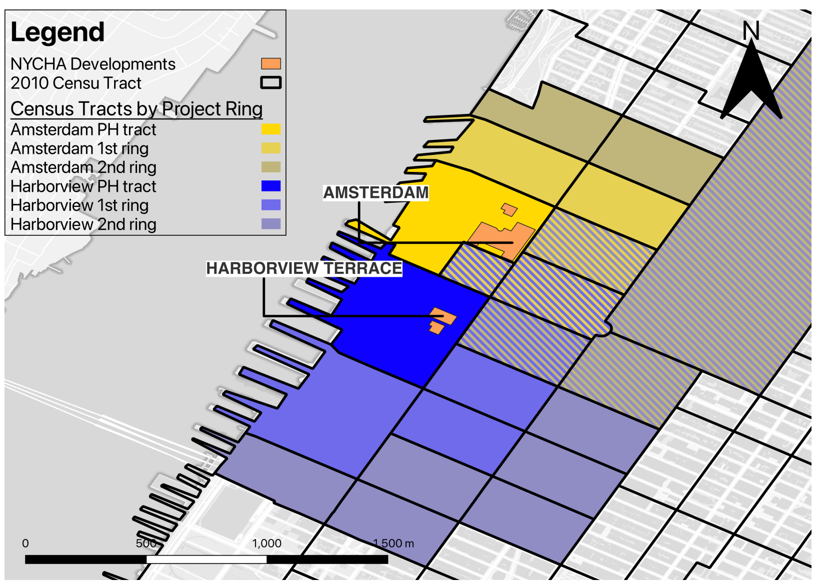
$$y_{k,m,p,t} = \sum_{R} \sum_{\tau=-60}^{60} \beta_{\tau,r} \cdot \mathbb{1}\left(t - Y_p = \tau, R = R(k,p)\right) + \mathbf{X}_{p,k} + \rho_{p,t} + \rho_{p,m,t} + u_{k,m,p,t} \quad (2)$$

The coefficient of interest, denoted as $\beta_{\tau,r}$, captures the effect of the arrival of public housing on demographics over time in each treated tract, relative to tracts in the outermost rings. I interact each time dummy with an indicator for the ring $R(k,p)$ in which a tract k around project p is located. Y_p denotes the year when a project p was completed and the set of rings is defined as $R = \{Treated, 1st\ ring\}$.

Project-specific controls are included to account for local infrastructure developments that may independently influence neighborhood outcomes ($\mathbf{X}_{p,k}$). Specifically, I control for whether a tract was exposed to (i) an urban renewal project, (ii) an arterial road expansion, or (iii) a new subway connection, using time-varying treatment indicators. These indicators are fully interacted by project (sub-experiment) ID to allow the timing and effects of these developments to vary flexibly across projects. In each sub-experiment, I also control for the baseline value of the outcome variable to account for pre-existing differences across rings prior to treatment.

In addition, I include project-by-year fixed effects ($\rho_{p,t}$) to capture time-specific shocks within each project. To further control for differential trends across space, I include neighborhood-by-year fixed effects ($\rho_{p,m,t}$) fully interacted with the project ID. This ensures that time trends are allowed to vary flexibly across neighborhoods within each project. All controls and fixed effects are thus implemented within sub-experiment strata, ensuring that identification relies solely on within-project comparisons over time.

Figure 1: Treatment construction



Note. Figure 1 provides an illustrative example of overlapping neighborhood/distance rings for two public housing projects: Harborview Terrace and Amsterdam Houses. The concept of neighborhood rings is depicted, with blue and yellow hatched census tracts representing the areas that belong to the respective public housing tract and are located within their respective rings. It is important to note that this tract may appear multiple times in the dataset. If a public housing tract was lying within a neighborhood ring to another public housing tract, it was excluded from the respective ring such that no treated tract appears in the control group.

This approach is equivalent to estimating Equation 2 separately for each project, then aggregating the coefficients using regression weights.[14] Since the number of treated and

[14]A stacked difference-in-differences design effectively accounts for heterogeneous treatment effects, a

control observations varies across sub-experiments, estimates are weighted by the relative frequency of tracts within each sub-experiment to ensure proportional representation. Standard errors are clustered at the sub-experiment level, which, in this case, corresponds to the project.[15]

However, this estimation strategy does not take into account general equilibrium effects, where projects could impact rents and population across the city. Additionally, projects can increase the supply of low-income housing in the city. Thus, I assume that such effects are minimal, with the most significant impact being concentrated near the projects. One concern is that individuals may relocate to nearby areas, which would violate the Stable Unit Treatment Value Assumption (SUTVA). In Appendix C in Figure C1.2, I show the deviation of the primary outcome variables, by treatment and control group, from the long-run trend of the average tract in the rest of New York City. The treatment group deviates substantially from the rest of New York City over time, while the control group closely follows the overall city trend. If individuals sorted themselves into the control areas, we would expect those areas to deviate from the citywide trend. If significant city-wide effects exist, my estimates may understate the full impact, but the relative comparisons across rings would remain unaffected. Additionally, since rent prices are forward-looking, effects may begin when construction is first announced. These anticipation effects are absorbed, as treatment effects are averages of all projects completed at any time within a census decade, and estimates are a composite of anticipation and completion effects.

5 Reduced Form Results

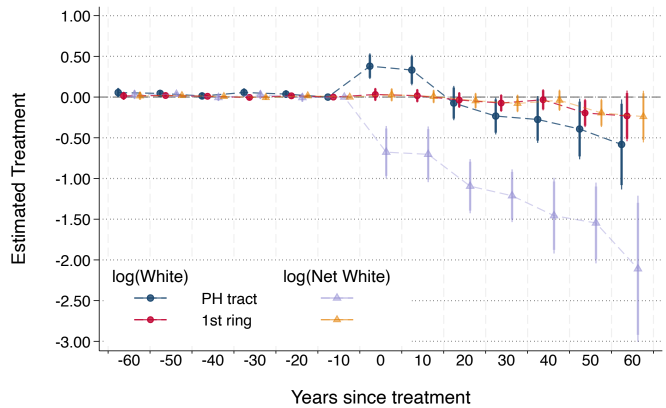
This section presents the empirical findings on the long-term effects of public housing on neighborhood composition. Using a difference-in-differences (DiD) framework, I estimate how the arrival of public housing projects affected the four equilibrium variables: Black population, White population, Hispanic population and rent prices. The analysis proceeds in two parts. First, I examine effects within a window of 60 years before and after construction on population outcomes. Second, I estimate the effect on rent prices comparing the census rent to the rental listings. Furthermore, I assess the robustness of these results using quality-adjusted rents, a subset of projects with non-overlapping rings, and a panel DiD framework. This approach enables the use of estimators that are robust to heterogeneous treatment effects, in contrast to the stacked DiD setup used in the main specification.

known limitation of traditional DiD estimators (Wang et al., 2024).

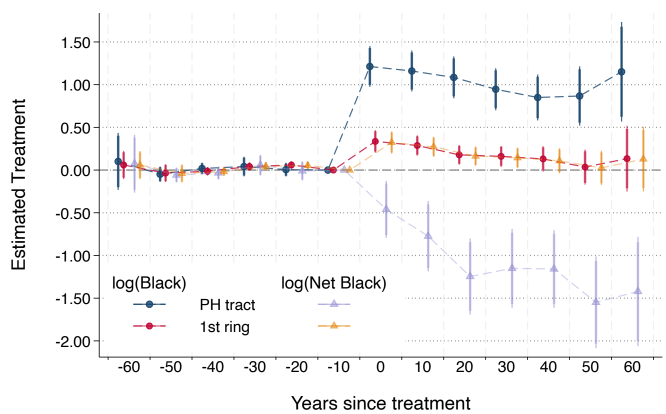
[15] In Appendix D4, I estimate a dynamic TWFE specification in a panel setup at the census tract level. I report event study results using the standard TWFE model along with the de Chaisemartin and D'Haultfoeuille estimator (De Chaisemartin and D'Haultfœuille, 2020).

Figure 2: Long-Run Effects on Demographics

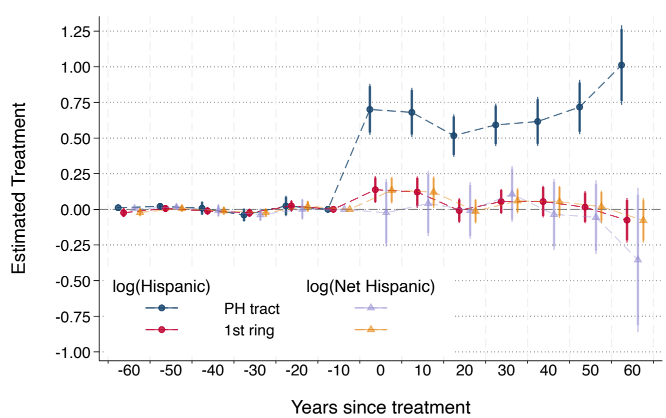
(a) White Population



(b) Black Population



(c) Hispanic Population



Note: Figure 2 plots the coefficients $\hat{\beta}_{\tau,r}$ from Equation 2. Each panel reports dynamic treatment effects on the log population of White, Black, and Hispanic residents. Blue and red lines indicate effects in treated and adjacent tracts, respectively. Purple and orange lines show analogous effects net of public housing residents. The omitted category comprises second-ring tracts. Standard errors are clustered at the project level. Vertical bars indicate 90% and 95% confidence intervals. Estimates are weighted by the number of census tracts in each ring.

Long-Run Population Effects. Using the full sample of public housing projects constructed before 1960, Figure 2 shows dynamic treatment effects up to 60 years after construction.[16] I separately estimate the effects on White, Black, and Hispanic population, as well as the population net of public housing residents.

Panel a shows that White population in treated tracts initially rises by 46.2% at the time of construction, followed by a steady decline to 44.1% below baseline after 60 years. These shifts are substantial given a baseline population of 3,410: White population increases by an average of 1,577 immediately after construction and declines by 1,504 after 60 years. In adjacent tracts, there is some evidence of spillovers, with declines of 17.7% and 20.6% observed 50 to 60 years post-construction. When excluding public housing residents (purple line), net White population in treated tracts drops by 49.2% immediately after construction and falls to –87.9% after 60 years. This indicates that private White residents were replaced by public housing tenants, and the long-run decline is driven by the private sector rather than movements within public housing.

Panel b displays a dramatic and persistent increase in Black population in treated tracts—rising by 236.4% immediately and stabilizing around 160% above baseline over six decades. With an average baseline of 1,503 Black residents, this corresponds to an initial increase of 3,553 and a net gain of 2,405 after 60 years. Adjacent tracts also show substantial increases of 14% to 40% in the first four decades, suggesting that the demographic effects of public housing extended beyond the project boundaries. However, when excluding public housing residents, the Black population in treated tracts declines by 36.9% immediately and by 75.9% after 60 years. This implies that the observed growth in total Black population is driven entirely by the influx of public housing tenants, which more than offsets the decline in privately housed Black residents.

Panel c presents the long-run effects on the Hispanic population. In treated tracts, Hispanic population rises sharply—by 101.6% in the first decade—followed by a continued increase to 175.1% sixty years after construction. This translates to a net gain of 1,316 residents, from a baseline of 752. Spillover effects into adjacent tracts are limited: modest increases of 14.8% to 12.9% in the first two decades taper off to near zero in the long run. Crucially, there is no discernible change in the Hispanic population excluding public housing residents, indicating that these gains are entirely attributable to direct placement into public housing. Unlike the patterns observed for White and Black populations, where private households responded strongly, the demographic changes for Hispanic residents appear to reflect administrative allocation rather than broader neighborhood sorting dynamics.

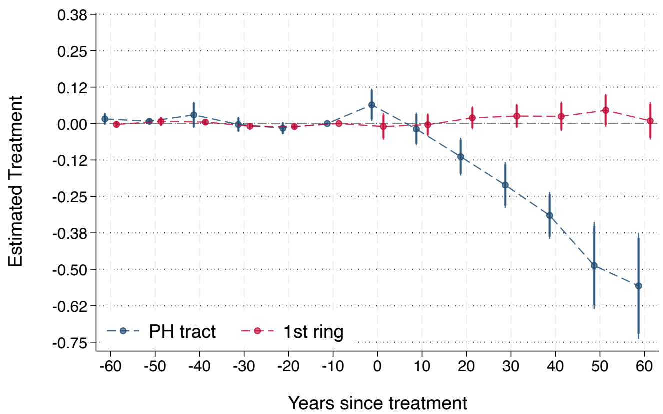
Together, these results highlight that public housing had large and durable effects on

[16] Because I find substantial effect sizes, I convert all point estimates from log points to percent using $exp(\hat{\beta}) - 1$.

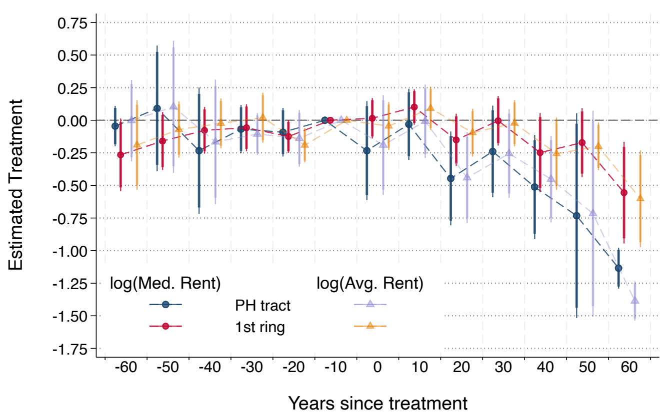
racial composition, suggesting that public housing triggered broader racial re-sorting of private households across nearby neighborhoods.

Figure 3: Long-Run Effects on Rents

(a) log(Census Rent)



(b) NYT Rent



Note: Figure 3 plots coefficients $\hat{\beta}_{\tau,r}$ from Equation 2. Panel a shows effects on the log of median contract rent from the U.S. Census. Panel b uses median and average rent per room at the census tract level from the New York Times (NYT) rental listings. Blue and purple lines indicate effects in treated and red and orange lines in adjacent tracts. The regression in Panel b excludes baseline controls, as not all tracts have consistent NYT coverage. The omitted category consists of tracts in the second ring. Standard errors are clustered at the project level. Vertical lines show 90% and 95% confidence intervals. Estimates are weighted by the number of census tracts within each ring.

Long-Run Rent Effects. I plot the effect on rent prices in Figure 3, using two complementary data sources. Panel a plots the impact on log median contract rents from the U.S. Census, while Panel b uses the average and median rent per room based on New York Times (NYT) rental listings. In both panels, treated census tracts (PH tracts) are compared to adjacent tracts (1st ring), with the second ring omitted as the reference group.

Panel a documents the long-run trajectory of census-based rents in treated tracts. Rents initially rise by 6.6% following construction, but this effect quickly reverses. Over the next three to six decades, rents fall substantially — by 19.0% to 42.7%—relative to baseline levels. In contrast, rents in adjacent tracts remain flat throughout, with no statistically significant deviations from pre-construction trends. These patterns suggest that the negative externalities associated with public housing—such as stigma, physical deterioration, or declining neighborhood demand—are highly localized and do not spill over into surrounding areas.

Panel b corroborates the long-run decline in rental values using listing data from the New York Times. The outcome variables are the median and average rent per room, constructed from a hedonic regression of listing-level rents. In treated tracts, rents fall by 21.3% (median) and 22.6% (average) after 30 years, with even steeper declines — 67.8% and 75.0%, respectively — observed after 60 years. While rent effects in adjacent areas are generally small and statistically insignificant, there is suggestive evidence of long-run spillovers, with average and median rents declining by 42.6% and 45.2%, respectively, six decades after construction.

The similarity in trends across both panels—despite differences in data construction, coverage, and measurement—provides strong evidence that public housing construction led to persistent and substantial rent reductions in the areas where it was built. The sharper declines in rental listings, which are more responsive to changes in neighborhood quality and demand, reinforce the interpretation that these effects reflect genuine market responses rather than administrative or compositional changes alone. Crucially, the concurrent decline in rents and the observed out-migration of privately housed residents suggest that public housing projects altered how neighborhoods were valued in the private market. Rather than affecting only the tenants placed in public units, these developments reshaped broader housing demand and perceptions of neighborhood quality over the long run.

5.1 Robustness Checks

Quality adjusted rent. A potential concern with using raw average or median rents from newspaper listings is the risk of compositional bias over time. For instance, if higher-quality listings become overrepresented in later years, this could mask underlying rent declines. To address this, I construct a quality-adjusted rent measure using a hedonic regression, as detailed in Appendix D1, taking quality differentials of listings and seasonality into account. Panel a in Figure D1.1 presents the estimated dynamic effects using this quality-adjusted rent. The estimated trajectory mirrors the census-based findings, with long-run declines in treated tracts—albeit slightly smaller in magnitude. For example, by 40 to 60 years after construction, the quality-adjusted rents fall by 18.4% to 36.4%, a range between the NYT rent per room estimates and the census median contract rent. However, the standard errors are larger, reflecting the fact that the hedonic approach is based on a smaller sample of listings and does not span all tracts or years. In summary, the consistency across rent measures supports the conclusion that public housing projects had large, localized, and persistent negative effects on surrounding rental markets—robust to quality adjustment and data source.

No overlapping Rings. To address potential biases arising from spatial spillovers between treated and control areas, I re-estimate the main event study specification using only a subset of 18 projects with non-overlapping 1st and control rings. In Appendix D3, I estimate the main event study specification using these 18 projects. This cleaner identification strategy addresses concerns about spatial interference across rings. Results from this exercise are shown in Figure D3.1. This robustness check confirms that the main findings are not driven by spatial overlap between treatment and control areas. While there are some differences in levels—likely due to sample composition — the long-run trends in neighborhood composition and rents closely track those from the full sample. This reinforces the conclusion that public housing had persistent, localized effects, and that the main estimates are not sensitive to spatial contamination across rings.

Treatment Heterogeneity. Finally, in Appendix D4 I use a standard panel DiD setting, benchmarking the Two-Way Fixed Effects (TWFE) estimator against the estimator proposed by De Chaisemartin and D'Haultfœuille (2020). Both TWFE and DCDH estimators closely match the stacked DiD estimates. It should be noted that the control group in the panel DiD setting is not as well defined as in the stacked version, where each project-tract is compared to a fixed set of control tracts. Rather, the comparison group includes all tracts that were never treated. Nevertheless, the results are consistent in magnitude with the stacked DiD estimates. Moreover, the TWFE estimates exhibit pre-trends, which can be explained by effect heterogeneity further discussed in Section 6.

6 Heterogeneity

An important question is whether the impact of public housing is uniform across neighborhoods or varies depending on local conditions. In this section, I explore heterogeneity in treatment effects along three key dimensions: construction period, borough, and neighborhood income.

Heterogeneity by Construction Period. A key question is whether the effects of public housing projects varied over time, particularly given the significant policy shifts in the 1970s. To address this, I split the sample at 1970 and estimate the model separately for projects built before and after this cutoff. The year 1970 marks a turning point in federal housing policy, as discussed in Section 2. Before 1970, public housing aimed to serve a mixed-income population and was often integrated into broader urban renewal plans. After 1970, housing policy shifted toward serving very low-income households, often in more isolated or stigmatized developments, with reduced federal investment and weakened enforcement of quality standards. Moreover, New York City was facing a severe fiscal crisis in the 1970s, leading to significant cutbacks in public services and infrastructure, which likely influenced the ability of public housing projects and their effects on neighborhoods. There are 166 projects built between 1930 and 1969 and 66 projects between 1970 and 2000.

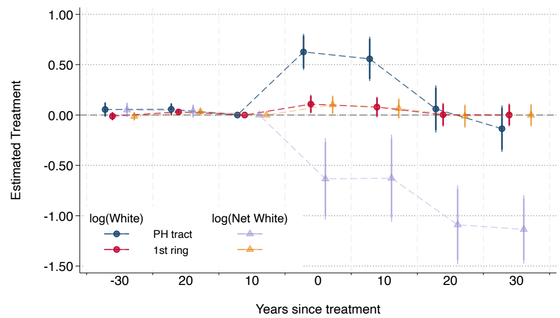
Figure 4 compares dynamic treatment effects on White, Black and Hispanic population, as well as median contract rents, for public housing projects completed before and after 1970, over a 30-year post-construction horizon. Panel a reveals that early public housing projects were initially accompanied by a sharp influx of White residents—an 87.2% increase relative to baseline (3843) in the year of construction, corresponding to a net gain of 3,351 individuals. However, this initial increase is not sustained. Over time, the White population in treated tracts declines steadily, culminating in a 12.7% drop by year 30—a loss of approximately 489 residents relative to baseline. Adjacent tracts see only a modest and short-lived increase of 11.4%, which quickly dissipates and becomes statistically indistinguishable from zero.

In contrast, post-1970 projects exhibit a markedly different pattern. Panel b shows that the White population in treated areas declines almost immediately—falling by 19.9% within the first decade and stabilizing around –25% in the longer run, implying a total loss of approximately 581 individuals. Once public-housing residents are excluded, the decline becomes even more pronounced: the net loss of White residents reaches 2,609 in pre-1970 tracts and 967 in post-1970 tracts. This suggests that post-1970 effects are largely driven by the outflow of non-public-housing White residents rather than changes in public housing assignment.

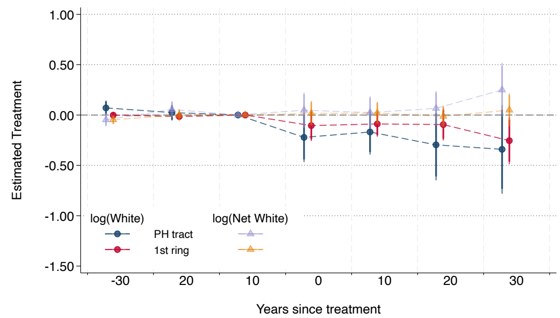
Taken together, these results suggest a clear shift in neighborhood dynamics associated with public housing over time. Early projects may have initially attracted some White residents, but ultimately catalyzed large-scale departures of non-public-housing White households. In contrast, post-1970 projects deterred in-migration from the outset, resulting in a more modest—yet persistent—decline in the White population. This contrast indicates that while earlier projects reshaped neighborhoods through gradual demographic turnover, later projects altered selection patterns at the point of entry.

Figure 4: Effect on Demographics by Construction Cohort

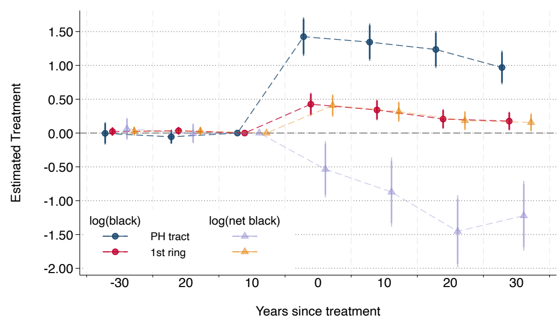
(a) White Population - pre 1970



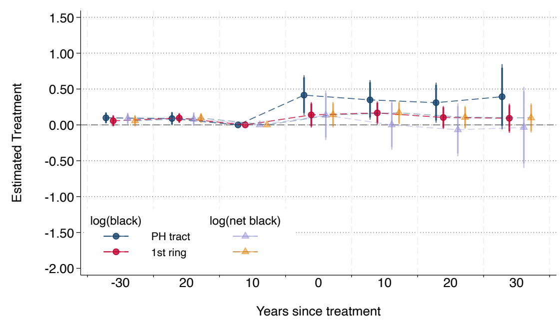
(b) White Population - post 1970



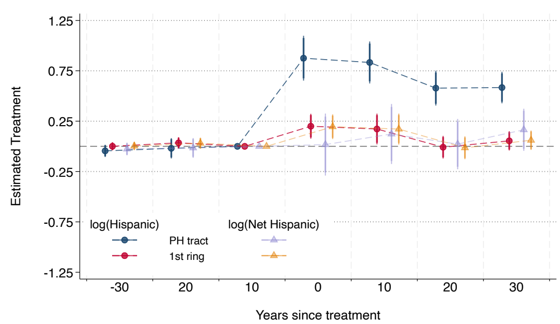
(c) Black Population - pre 1970



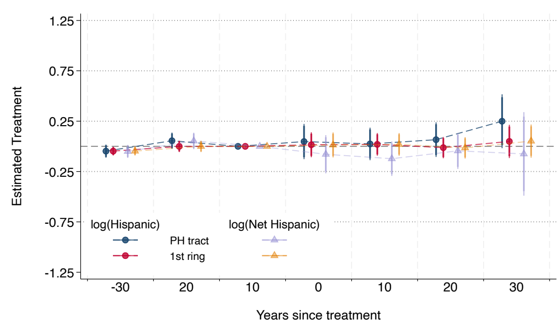
(d) Black Population - post 1970



(e) Hispanic Population - pre 1970



(f) Hispanic Population - post 1970



Note: Figure 4 plots dynamic treatment effects $\hat{\beta}_{\tau,r}$ from Equation 2, estimated separately for public housing projects completed before and after 1970. Panels a,c, and e show effects for pre-1970 projects; Panels b,d, and f display results for post-1970 projects. Each panel reports treatment effects on the log population of White, Black, and Hispanic residents. Blue and red lines indicate estimates for treated and adjacent tracts, respectively. Purple and orange lines represent corresponding effects net of public housing residents. The omitted category is census tracts in the second ring. Standard errors are clustered at the project level. Vertical bars denote 90% and 95% confidence intervals. Estimates are weighted by the number of census tracts within each ring.

Panels c and d highlight stark differences in Black population dynamics before and after 1970. The Black population in treated tracts increased by 315.79% during the first decade—an average rise of 4,592 individuals from a baseline of 1,454. Adjacent areas also experienced long-run gains of 53.27%, or 519 individuals compared to a baseline Black population of 975. However, the growth slows over time, settling at 163.53% in treated tracts and 19.36% in adjacent areas 30 years post-construction. Importantly, the number of Black residents living in private housing in treated areas declined substantially—from –41.43% in the first decade to –70.54% by year 30— corresponding to a loss of approximately 1,026 Black residents in private housing.

By contrast, post-1970 projects produced smaller but more stable changes. The total Black population in treated tracts rose by approximately 45%, or 732 individuals, with no significant spillovers to adjacent areas. Notably, the private Black population remained unaffected, indicating that the observed increase was primarily driven by public housing allocations rather than broader shifts in neighborhood composition.

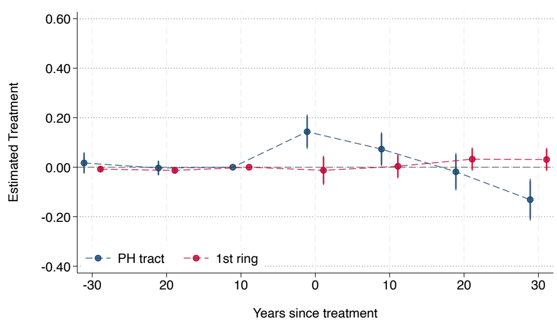
Unlike the patterns observed for White and Black populations, the dynamics for Hispanic residents differ in two key ways. First, as shown in Panels e and f, the total Hispanic population in treated tracts for pre-1970 projects increased by 140.13%, or an average of 661 individuals from a baseline of 472—substantial but more moderate than the effects observed for Black residents. Second, the private Hispanic population remains statistically unchanged, indicating that the observed increase is almost entirely attributable to public housing assignments. Moreover, projects constructed after 1970 exhibit no significant effect on either total or private Hispanic population, suggesting limited demographic response beyond direct placement into public housing.

In Figure 5, I examine how rents evolved following public housing construction. For projects completed before 1970, median contract rents in treated tracts initially rose by 15.37% but then declined steadily, reaching –12.28% thirty years after completion (Panel a). By contrast, post-1970 projects show no evidence of a short-term increase and instead exhibit persistent rent declines, with effects as large as –23.28% three decades later (Panel b). Notably, when focusing on residents excluding those in public housing, the effects are essentially zero. This suggests that later demographic changes were driven not by broader market responses, but by the direct placement of households into public housing.

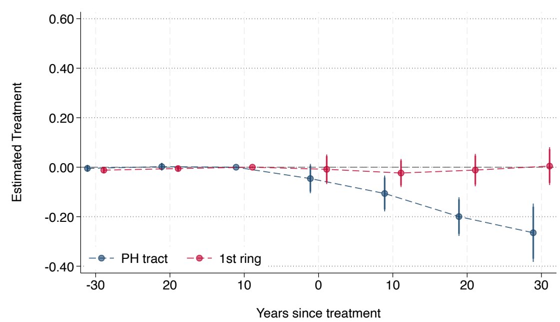
Using the data on median and average rent per room from New York Times listings, Panel c confirms a consistent pattern of negative rent effects. For projects constructed before 1970, rents initially rise modestly—by 6.29% (median) and 16.77% (average) — but fall sharply by approximately 29% after 20 years, before partially rebounding to –19% after 30 years. In contrast, later projects trigger an immediate decline in rents of about –60% which stabilizes around –39% after three decades. However, these effects are imprecisely estimated,

Figure 5: Effect on Rents by Construction Cohort

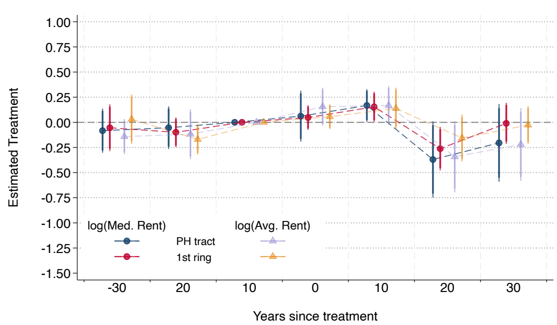
(a) log(Census Rent) - pre 1970



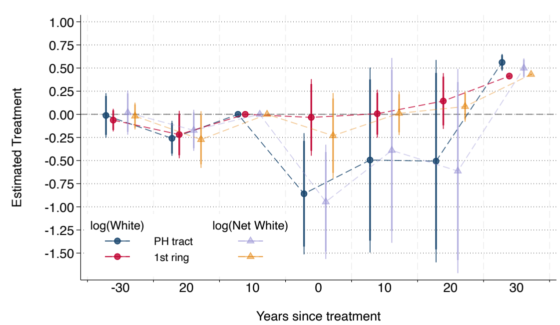
(b) log(Census Rent) - post 1970



(c) NYT Rent - pre 1970



(d) NYT Rent - post 1970



Note: Figure 5 reports dynamic treatment effects $\hat{\beta}_{\tau,r}$ from Equation 2, estimated separately for public housing projects constructed before and after 1970. Panels a and c display results for projects constructed before 1970, using U.S. Census median contract rent and New York Times (NYT) rental listings (median rent per room), respectively. Panels b and d show corresponding results for projects built after 1970, using Census and NYT average rent per room data. Blue and purple lines indicate effects in treated tracts; red and orange lines represent adjacent tracts. Regressions in Panels c and d omit baseline controls due to limited NYT coverage. The omitted category is tracts in the second ring. Standard errors are clustered at the project level. Vertical bars denote 90% and 95% confidence intervals. All estimates are weighted by the number of census tracts within each ring.

in part due to shifts in the composition of advertised properties. Figure D1.1c illustrates that adjusting for observable property characteristics yields a more moderate—but statistically insignificant—decline in rents. This is consistent with the absence of strong behavioral responses among private market renters, suggesting limited or muted market responses.

Overall, these findings reinforce the conclusion that early public housing construction had a lasting impact on neighborhood racial composition, primarily through behavioral responses by privately housed residents. These effects appear to drive the reduced-form results shown in Figure 2. In contrast, projects built after 1970 had more limited impacts, with changes largely confined to the direct placement of public housing residents, and little evidence of broader neighborhood adjustment. This contrast reflects a fundamental shift in the design and consequences of public housing policy after 1970. These results are consistent with prior evidence from New York City showing that federal public housing constructed between 1977 and 2000 did not reduce nearby property values (Ellen et al., 2007).

Heterogeneity by Borough I examine potential heterogeneity in treatment effects across New York City's five boroughs: Manhattan, Brooklyn, Queens, the Bronx, and Staten Island. Each borough represents a distinct geographic and socioeconomic environment. Queens and Staten Island are primarily suburban in character, with lower population density and more single-family homes, whereas Manhattan, Brooklyn, and the Bronx are more urbanized and densely built. The boroughs also differ substantially in their historical development patterns, industrial composition, and demographic structure.

To test for heterogeneity, I interact the event-time indicators with borough dummies, allowing treatment effects to vary across boroughs. The results, shown in Section D2, Figure D2.1, suggest that while the magnitude of effects differs somewhat across boroughs, the overall patterns are qualitatively consistent. In all five boroughs, public housing is associated with long-run demographic shifts and declines in rent—indicating broadly similar sorting dynamics despite institutional and spatial differences.

Heterogeneity by Neighborhood Income. Second, I explore heterogeneity by neighborhood income level, using pre-treatment income tertiles defined at the neighborhood (NTA) level. As shown by Diamond and McQuade (2019), affordable housing can signal neighborhood improvement in low-income areas, but may depress house prices in higher-income neighborhoods. To test this, I classify neighborhoods into income tertiles based on baseline median household income, deflated to 2010 dollars using the Consumer Price Index. I then interact the treatment indicator with indicators for each income tertile and present the results in Section D2, Figure D2.2. The trajectories of demographic composition and rents are strikingly similar across income groups, with only modest variation in magnitude. This suggests that public housing triggered similar sorting responses regardless of local income level.

7 Location Choice Model

In this section, I present a theoretical sorting model following Bayer, Ferreira, et al. (2007) and Almagro et al. (2023) to formalize how public housing influences individual preferences over neighborhoods. The model is designed to capture the response of the private housing market to public housing construction—specifically, how privately housed residents adjust their location choices in response to public housing presence and design.[17]

To link the model to the empirical analysis, I derive comparative statics that map changes in public housing presence to equilibrium outcomes in rents and population composition. These comparative statics allow me to decompose the reduced-form treatment effects estimated in Section 4 into underlying preference parameters. By doing so, I quantify how much different groups value or disvalue public housing in their neighborhood, providing a structural interpretation of the observed effects. The full model is derived in Appendix A.

The city consists of a total population $N = \sum_g N^g$, where residents belong to one of three ethnic groups $g \in \{B, W, H\}$, representing Black, White, and Hispanic households, respectively. Each resident i derives utility from consuming a single, city-wide final good, housing services, neighborhood amenities, and an idiosyncratic location-specific preference shock. The log indirect utility for a resident i of group g living in census tract k is given by:

$$V_{ik}^g = \varphi_{ik}^g + \epsilon_{ik} \tag{3}$$

where φ_{ik}^g is the component of indirect utility for census tract k that is common to all residents of group g, called mean indirect utility hereafter, and ϵ_{ik} is an idiosyncratic shock which is drawn from an extreme value type I distribution. The common component of indirect utility is:

$$\varphi_{ik}^g = \beta_{PH_1}^g PH_{1k} + \beta_{PH_2}^g PH_{2k} + \log(w_k) - \alpha \log(r_k) \tag{4}$$

Here PH_{Rk} is an indicator variable for tract k containing a public housing project or

[17] I do not model the allocation of public housing residents explicitly. Instead, I treat their assignment as exogenous to the market, allowing me to isolate and analyze the preferences and movements of residents in market-rate housing. This simplifying assumption ensures that the model speaks directly to the behavioral responses that are observed in the empirical analysis—namely, how public housing affects the broader neighborhood rather than the internal functioning of the program.

being adjacent to one ($R \in \{PH\ tract, \text{1st ring}\}$). Wages w_k are determined competitively and r_k are rents in census tract k. The model is closed by assuming an isoelastic supply function such that the number of housing units in tract k is given by:

$$S_k = \delta_k r_k^{\phi_k} \tag{5}$$

where δ_k is a supply shifter and ϕ_k is the tract-specific supply elasticity. Details of the model equilibrium are provided in Appendix A. Next, I derive two comparative statics to evaluate the equilibrium response of rent prices and population to public housing. First I differentiate the equilibrium rents with respect to PH, which yields:

$$\frac{d\log(r_k^*)}{dPH_{Rk}} = \frac{\Xi_r^g}{\phi_k + \alpha} \tag{6}$$

Equation 6 shows that equilibrium rent changes in response to public housing depend on three key factors: the income share spent on housing (α), the local housing supply elasticity (ϕ_k), and the utility-weighted average preference for public housing across all ethnic groups, denoted by Ξ_R^g. The more positively residents value public housing (higher Ξ_R^g), or the less elastic the housing supply (lower ϕ_k), the more rents rise. Using Equation 6 gives the equilibrium population response of group g in tract k to the construction of public housing:

$$\frac{d\log(N_k^{g*})}{dPH_{Rk}} = \beta_{PH_R}^g - \alpha \cdot \frac{\Xi_R^g}{\phi_k + \alpha} \tag{7}$$

Equation 7 characterizes how the population of group g responds to the introduction of public housing. The total response combines the group-specific preference for public housing ($\beta_{PH_R}^g$) and an adjustment for the rent increase, scaled by the income share of housing and the local supply elasticity. Intuitively, if public housing increases rents, some households will be priced out, attenuating the population response. Only in the absence of any rent effect (i.e., when $\partial \log(r_k^*)/\partial PH_{Rk} = 0$) does $\beta_{PH_R}^g$ reflect a pure preference effect.

To recover these preference parameters, I use reduced-form estimates of the effects of public housing on population and rents. Since the model's comparative statics directly map to the difference-in-differences estimates, I compute the preference parameters using the decomposition implied by Equation 7 and Equation 6. This approach enables a structural interpretation of the empirical results and links observed demographic shifts to underlying economic preferences:

$$\begin{aligned} \beta_{PH_R,t}^g &= \frac{\partial \log N_{kt}^{g*}}{\partial PH_{R,t}} + \alpha \times \frac{\partial \log r_{kt}^*}{\partial PH_{R,t}} \\ &= \hat{\beta}_{PH_R,t}^g + \alpha \times \hat{\beta}_{PH_R,t}^{rent} \end{aligned} \tag{8}$$

Here, $\hat{\beta}^{g}_{PH_R,t}$ is the estimated reduced-form effect of public housing on the log population of ethnic group $g \in \{W, B, H\}$, and $\hat{\beta}^{rent}_{PH_R,t}$ is the corresponding reduced-form effect on rental prices. Adding time subscripts, I distinguish between medium-run (t = 0–30 years) and long-run (t = 40–60 years) impacts. I fix the housing expenditure share at $\alpha = 0.3$ for all computations. To assess the precision of these preference estimates, I calculate their sampling variance based on the variances of the underlying reduced-form coefficients.

$$\mathrm{Var}(\beta^{g}_{PH_R,t}) = \mathrm{Var}(\hat{\beta}^{g}_{PH_R,t}) + \alpha^2 \mathrm{Var}(\hat{\beta}^{rent}_{PH_R,t}) \tag{9}$$

I provide a full derivation of this decomposition and variance expression in Appendix A1.

8 Public Housing Characteristics

In this section, I apply the framework developed in Section 7 to quantify how private (non–public housing) residents value different attributes of nearby public housing. I begin by estimating the effects of key, time-invariant project characteristics—including building height, total number of apartments, construction cost per room, ground coverage, and the initial racial composition of tenants—on four equilibrium outcomes: the populations of White, Black, and Hispanic residents not living in public housing, and the median contract rent in the census tract.[18]

Next, I use these reduced-form estimates to recover households' marginal willingness to pay (MWTP) for each public housing attribute. As shown in Section 7, MWTP can be derived through a decomposition of preference parameters, allowing me to estimate how much private residents are willing to pay in rent to live near—or avoid—particular types of public housing. I present results separately for the medium run (0–30 years post-construction) and the long run (40–60 years).

The analysis focuses on projects built before 1970, where the assumptions of the equilibrium sorting model — especially market-driven relocation — are more likely to hold. As discussed in Section 6, projects constructed after 1970 produced more muted neighborhood changes that appear largely driven by administrative placement of public housing residents, rather than by market-based responses. Nevertheless, I report MWTP estimates for post-1970 projects in Appendix D5.

[18] I use the census-based median contract rent for two reasons. First, it captures the central market rent trend relevant for identifying shifts in residential preferences. Second, listing-based rents from newspaper data suffer from limited and uneven spatial coverage, introducing substantial measurement error.

Public Housing Characteristics. To examine how specific design features of public housing shape neighborhood dynamics, I estimate the following regression for projects constructed before 1970:

$$y_{k,m,p,t} = \sum_{\substack{\tau \in \{0\text{–}30,\\ 40\text{–}60\}}} \sum_{i \in \mathbf{ATTR}} (\gamma_{0i\tau} \cdot PH\ tract_{p,t} + \gamma_{1i\tau} \cdot 1st\ ring_{p,t}) \times ATTR_i + \mathbf{X}_{p,k} + \rho_{p,t} + \rho_{p,m,t} + u_{k,m,p,t} \tag{10}$$

Here, **ATTR** denotes a vector of time-invariant public housing characteristics — each interacted with indicators for whether a tract is directly treated ($PH\ tract_{p,t}$) or lies in the first adjacent ring ($1st\ ring_{p,t}$). This setup allows me to separately estimate effects in the medium run (0–30 years post-construction) and long run (40–60 years post- construction), using only projects completed prior to 1970.

The vector **ATTR** includes five core attributes: (i) average construction cost per room[19], which serves as a proxy for building quality; (ii) the total number of public housing apartments in the tract, capturing the scale of the project; (iii) average project height, representing vertical density; (iv) average ground coverage, measured as the built-up area as a share of the parcel; and (v) the change in racial composition from the time of project opening to its composition in 2010, capturing shifts in tenant demographics.

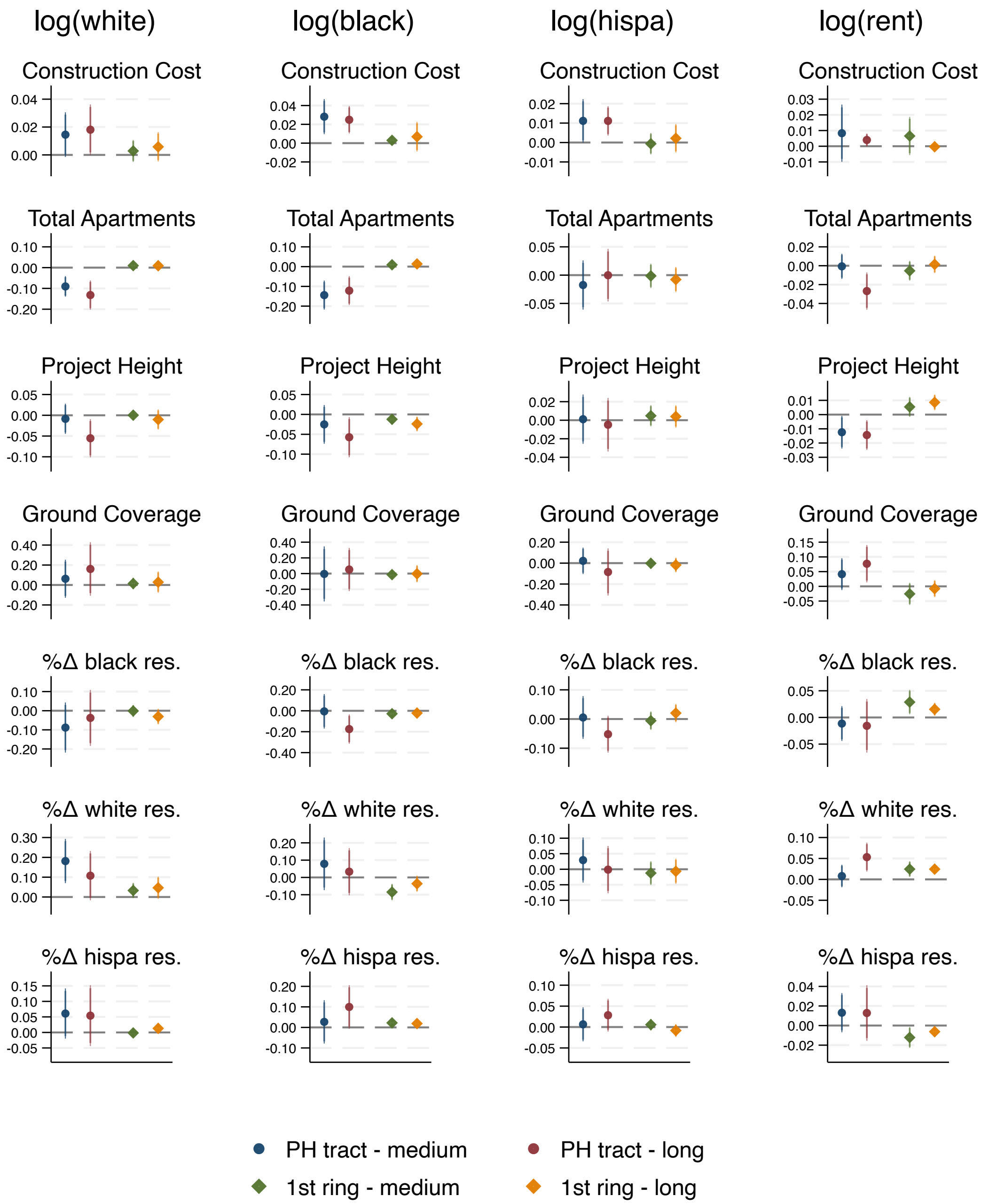
Each of these attributes addresses longstanding concerns about how public housing design and tenant composition shape neighborhood outcomes. The *Tower in the Park* model — characterized by tall buildings set within expansive green space — has been widely criticized for fostering social isolation, weakening community oversight, and increasing crime risks (Jacobs, 1992; Newman, 1997). Finally, changes in racial composition allow for testing household preferences for living near public housing with differing tenant demographics.

Figure 6 illustrates how key public housing characteristics influence neighborhood composition and rent levels, focusing specifically on private residents—that is, populations net of public housing tenants. Higher construction costs, interpreted as a proxy for building quality, are associated with modest but statistically significant increases in the net White, Black, and Hispanic populations in treated areas. A \$1,000 increase in construction costs corresponds to population gains of 1.5%, 2.8%, and 1.1% in the medium run and 1.8%, 2.5%, and 1.1% in the long run, respectively. Rents also rise slightly in treated tracts—by 0.8% in the medium run and 0.4% in the long run—while effects in adjacent tracts remain negligible.

In contrast, larger-scale developments—measured by the total number of public housing units—are associated with sizable population declines among non–public housing residents.

[19] This excludes any other development costs

Figure 6: Effects of Public Housing Attributes



Note: Figure 6 plots the coefficients $\hat{\gamma}_{0i\tau}$ and $\hat{\gamma}_{1i\tau}$ from Equation (10). The outcome variables are net White, Black, and Hispanic populations, defined as the total number of residents of each group in the tract minus the number residing in public housing. Median contract rent is also included as an outcome. Each treatment dummy—public housing tract and adjacent tract—is interacted with continuous public housing characteristics aggregated at the tract level. These characteristics include: (i) total number of public housing apartments, (ii) average building height, (iii) average ground coverage, (iv) average construction cost per room in 2010 dollars, and (v) the percentage point change in the racial composition of tenants at the project's initial opening. Standard errors are clustered at the project level; vertical bars denote 90% and 95% confidence intervals. All estimates are weighted by the frequency of census tracts within a ring; the omitted category consists of tracts in the second ring.

A 100-unit increase leads to medium-run declines of 9.0% (White) and 14.4% (Black), and long-run declines of 13.2% and 12.1%, respectively. These projects are also linked to a 2.7% decrease in median contract rents in the long run.

Building height plays a significant role in shaping long-run neighborhood composition. In treated tracts, a one-story increase in average project height reduces the net White and Black populations by 5.5% and 5.7%, respectively, over the long run. Rents also fall by 1.2% in the medium run and 1.4% in the long run. These patterns suggest that taller public housing structures are consistently associated with lower desirability among private residents. By contrast, responses among Hispanic households and rent effects in adjacent tracts are statistically negligible.

Ground coverage — defined as the share of parcel area occupied by built structures — offers a complementary perspective. While demographic responses to greater ground coverage are generally insignificant, the design attribute is associated with modest rent increases of 4.1% in the medium run and 7.7% in the long run in treated tracts. This finding supports the interpretation that compact, low-rise developments may be perceived as more desirable environments by the private market.

Changes in the racial composition of public housing residents, from project opening to 2010, yield nuanced effects on neighborhood demographics and rents. A 10 percentage point increase in the share of Black residents over this period is associated with long-run declines in the net Black (–17.5%) and Hispanic (–5.2%) populations, suggesting that even co-ethnic households may avoid areas with growing concentrations of Black public housing residents. Conversely, a 10 percentage point increase in the White resident share from opening to 2010 leads to an 18.1% rise in the net White population in the medium run, along with a 5.3% increase in median rents in treated tracts and 2.5% in adjacent areas. Changes in the Hispanic share over time yield weak and generally insignificant effects across all groups.

These results indicate that private households respond not only to physical features of public housing, but also to the evolving composition of its residents—highlighting the importance of perceived neighborhood trajectories over time. While higher building quality and compact designs are generally associated with increased desirability, larger and taller developments appear to deter private households. Importantly, variation in the racial composition of public housing tenants over time — particularly increases in the share of Black or White residents — produces divergent responses across racial groups. However, these estimates, while suggestive, do not provide a direct monetary interpretation. In the next section, I translate these effects into marginal willingness to pay (MWTP) estimates to quantify the rent premiums or discounts that households associate with different public housing characteristics.

Marginal Willingness to Pay. A central question in evaluating the long-run impacts of public housing is how much households value different characteristics of nearby developments. To address this, I estimate households' marginal willingness to pay (MWTP) for specific attributes of public housing projects. The MWTP for a given attribute i reflects the rent premium that a household is willing to pay for a marginal improvement (or requires in compensation for a deterioration) in that attribute, holding utility constant. Formally, and following Equation 3, it is defined as:

$$\text{MWTP}^g_{\text{ATTR}_{i,t}} = -\frac{\dfrac{\partial V}{\partial \text{ATTR}_{i,t}}}{\dfrac{\partial V}{\partial \log r_{kt}}} = \frac{\beta^g_{\text{ATTR}_{i,t}}}{\alpha} \,. \tag{11}$$

Here, $\beta^g_{\text{ATTR}_{i,t}}$ denotes the preference parameter for attribute i among households of group g. This parameter reflects the trade-off residents make between neighborhood amenities—captured by public housing characteristics—and housing costs. Because the derivative is taken with respect to $\log r_{mt}$, the MWTP is expressed in log-point terms for a unit change in the attribute. I convert these values into monthly dollar amounts using the average median contract rent by ring and time period. This yields a clear and policy-relevant interpretation: the dollar amount that privately housed residents are willing to pay, or require in compensation, to live near public housing with a given feature.

To recover these preference parameters, I combine the reduced-form treatment effects from earlier sections with the equilibrium relationships derived in Figure 6. In particular, I use the comparative statics that relate population and rent responses to changes in public housing attributes. This decomposition, outlined in Equation 8, allows me to isolate households' marginal valuations of specific public housing features across racial and ethnic groups. Finally, I assess the statistical precision of the resulting MWTP estimates by computing their sampling variances, based on the empirical variances of the underlying reduced-form coefficients. A full derivation of this decomposition and the variance formula is provided in Appendix A1.

The resulting MWTP estimates are reported in Table 1 through Table 3, separately for White, Black, and Hispanic residents not living in public housing. Estimates are presented for both the medium run (0–30 years after construction) and the long run (40–60 years), covering the key attributes discussed earlier.

Higher-cost, well-built public housing projects are consistently valued more by all racial and ethnic groups in the private housing market. A \$1,000 increase in construction costs – used as a proxy for building quality – is associated with statistically significant and positive MWTP values for White, Black, and Hispanic households not living in public housing. In treated tracts, White residents are willing to pay \$25.68 more per month in the medium

Table 1: MWTP Estimates for White Residents

	0-30 years		40-60 years	
	MWTP (%)	MWTP ($)	MWTP (%)	MWTP ($)
	PH Tract			
# Apartments	-0.30^{***}	-136.11	-0.47^{***}	-287.87
	(0.083)		(0.119)	
Construction cost	0.06^{**}	25.68	0.06^{**}	39.74
	(0.029)		(0.031)	
# Stories	-0.04	-18.35	-0.20^{**}	-122.74
	(0.062)		(0.078)	
%Δ Black res.	-0.31	-137.80	-0.14	-87.05
	(0.222)		(0.249)	
%Δ Hisp. res.	0.22	97.65	0.19	119.39
	(0.137)		(0.166)	
%Δ White res.	0.61^{***}	276.71	0.41^{*}	254.78
	(0.189)		(0.211)	
% Ground coverage	0.25	112.81	0.61	378.01
	(0.323)		(0.454)	
	1st Ring			
# Apartments	0.03	15.62	0.03	29.70
	(0.029)		(0.045)	
Construction cost	0.02	9.06	0.02	17.38
	(0.015)		(0.018)	
# Stories	0.01	3.64	-0.03	-23.40
	(0.021)		(0.041)	
%Δ Black res.	0.02	13.92	-0.09	-79.07
	(0.054)		(0.069)	
%Δ Hisp. res.	-0.02	-9.87	0.04	34.18
	(0.024)		(0.030)	
%Δ White res.	0.13^{**}	75.73	0.18^{*}	164.55
	(0.066)		(0.095)	
% Ground coverage	0.02	13.15	0.09	79.59
	(0.101)		(0.181)	

Note: This table reports marginal willingness to pay (MWTP) estimates for public housing characteristics for private White residents — that is, residents not living in public housing — derived from Equation 10 using coefficients from Figure 6. Estimates are shown for two post-treatment horizons: the medium run (0–30 years) and the long run (40–60 years). MWTP values represent the percentage increase in monthly rent a household would be willing to pay for a one-unit change in the respective attribute (Columns 2 and 4). Changes are normalized as follows: 100 units for Apartments, $1,000 for Construction Costs, one storey for Height, and 10 percentage points for Ground Coverage and Racial Composition. Rent values are based on the average median contract rent (in 2010 dollars) by ring and time period (Columns 3 and 5). Each estimate is reported separately for public housing tracts (PH tracts) and adjacent areas (1st ring). The sample includes only projects constructed before 1970. Standard errors for the MWTP estimates are computed from the empirical variance of the underlying population and rent coefficients, as detailed in Appendix A1.
Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

run, rising to \$39.74 in the long run. Black residents show the strongest preferences, with MWTP estimates of \$46.20 and \$53.81, respectively. Hispanic residents exhibit smaller but still positive valuations of \$20.56 (medium run) and \$25.41 (long run).

Estimates for adjacent areas are generally positive but not statistically significant, suggesting that the perceived value of high-quality construction is concentrated within the directly treated neighborhood. Overall, these results highlight that higher construction quality is broadly appreciated, particularly by Black households, and contributes to neighborhood desirability.

MWTP estimates for public housing apartments reveal consistently negative valuations among private White and Black residents. A 100-unit increase in public housing supply is associated with statistically significant decreases in willingness to pay: White residents would require monthly compensation of \$136.11 in the medium run and \$287.87 in the long run to remain in treated tracts, while Black residents would require \$217.33 and \$266.76, respectively. Although estimates for adjacent areas follow a similar pattern, they are generally smaller in magnitude and not statistically significant. Hispanic households also exhibit negative MWTP values, but these effects are not statistically distinguishable from zero. Overall, these results suggest that larger public housing developments are perceived as less desirable by private residents, possibly due to concerns about density, neighborhood stigma, or associated declines in market rent levels.

The height of public housing buildings plays a significant role in shaping long-run neighborhood desirability. Among private residents in treated tracts, the marginal willingness to pay (MWTP) for a one-story increase in building height is consistently negative. For White households, MWTP declines from –\$18.35 in the medium run to –\$122.74 in the long run. For Black households, the drop is even steeper, from –\$43.39 to –\$126.93 over the same horizon. Hispanic households show smaller and statistically insignificant responses, suggesting weaker preferences regarding vertical density. In adjacent areas, MWTP estimates are generally imprecise. However, the long-run MWTP for Black households in adjacent tracts remains sizable and negative (–\$65.24), indicating that the aversion to high-rise designs extends beyond the immediate project footprint.

In contrast, ground coverage — defined as the share of the parcel occupied by built structures — shows no clear impact on demographic sorting. While MWTP estimates for Whites and Blacks suggest mild preferences for denser, low-rise designs, these effects are statistically insignificant. Hispanic households exhibit a small but statistically significant negative MWTP for increased ground coverage in treated tracts in the long run, implying some aversion to environments that are highly built up. Across all groups, MWTP estimates in adjacent tracts remain negligible.

Taken together, these findings lend support to long-standing critiques of the *Tower in*

Table 2: MWTP Estimates for Black Residents

	0-30 years		40-60 years	
Variable	MWTP (%)	MWTP ($)	MWTP (%)	MWTP ($)
		PH Tract		
# Apartments	-0.48^{***}	-217.33	-0.43^{***}	-266.76
	(0.124)		(0.124)	
Construction cost	0.10^{***}	46.20	0.09^{***}	53.81
	(0.033)		(0.025)	
# Stories	-0.10	-43.39	-0.21^{**}	-126.93
	(0.083)		(0.085)	
%Δ Black res.	-0.03	-13.17	-0.60^{**}	-369.63
	(0.279)		(0.237)	
%Δ Hisp. res.	0.10	46.71	0.35^{*}	213.26
	(0.179)		(0.179)	
%Δ White res.	0.27	122.12	0.16	101.53
	(0.257)		(0.231)	
% Ground coverage	0.03	12.16	0.25	154.67
	(0.593)		(0.466)	
		1^{st} Ring		
# Apartments	0.03	14.73	0.05	42.69
	(0.039)		(0.041)	
Construction cost	0.02	9.34	0.02	20.68
	(0.010)		(0.027)	
# Stories	-0.04	-20.07	-0.07^{**}	-65.24
	(0.023)		(0.032)	
%Δ Black res.	-0.06	-36.17	-0.06	-52.27
	(0.083)		(0.089)	
%Δ Hisp. res.	0.06^{*}	35.24	0.06	52.24
	(0.035)		(0.039)	
%Δ White res.	-0.26^{***}	-145.74	-0.10	-87.78
	(0.081)		(0.079)	
% Ground coverage	-0.07	-39.63	-0.01	-13.50
	(0.111)		(0.186)	

Note: This table reports marginal willingness to pay (MWTP) estimates for public housing characteristics for private Black residents — that is, residents not living in public housing — derived from Equation 10 using coefficients from Figure 6. Estimates are shown for two post-treatment horizons: the medium run (0–30 years) and the long run (40–60 years). MWTP values represent the percentage increase in monthly rent a household would be willing to pay for a one-unit change in the respective attribute (Columns 2 and 4). Changes are normalized as follows: 100 units for Apartments, $1,000 for Construction Costs, one storey for Height, and 10 percentage points for Ground Coverage and Racial Composition. Rent values are based on the average median contract rent (in 2010 dollars) by ring and time period (Columns 3 and 5). Each estimate is reported separately for public housing tracts (PH tracts) and adjacent areas (1st ring). The sample includes only projects constructed before 1970. Standard errors for the MWTP estimates are computed from the empirical variance of the underlying population and rent coefficients, as detailed in Appendix A1.
Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 3: MWTP Estimates for Hispanic Residents

	0-30 years		40-60 years	
	MWTP (%)	MWTP ($)	MWTP (%)	MWTP ($)
		PH Tract		
# Apartments	−0.06	−26.51	−0.03	−16.71
	(0.074)		(0.079)	
Construction cost	0.05**	20.56	0.04***	25.41
	(0.021)		(0.013)	
# Stories	−0.01	−3.91	−0.03	−18.83
	(0.045)		(0.049)	
%Δ Black res.	0.01	2.93	−0.19*	−116.41
	(0.125)		(0.110)	
%Δ Hisp. res.	0.03	15.79	0.11	66.32
	(0.070)		(0.067)	
%Δ White res.	0.11	48.17	0.05	30.57
	(0.123)		(0.130)	
% Ground coverage	0.11	51.27	−0.21	−127.57
	(0.215)		(0.382)	
		1^{st} Ring		
# Apartments	−0.01	−5.48	−0.02	−22.19
	(0.037)		(0.038)	
Construction cost	0.00	2.53	0.01	6.37
	(0.011)		(0.013)	
# Stories	0.02	12.16	0.02	20.17
	(0.020)		(0.021)	
%Δ Black res.	0.01	5.72	0.08	75.94
	(0.055)		(0.053)	
%Δ Hisp. res.	0.01	3.89	−0.03	−30.86
	(0.024)		(0.026)	
%Δ White res.	−0.02	−9.38	0.00	2.59
	(0.065)		(0.069)	
% Ground coverage	−0.03	−17.71	−0.07	−61.88
	(0.091)		(0.118)	

Note: This table reports marginal willingness to pay (MWTP) estimates for public housing characteristics for private Hispanic residents — that is, residents not living in public housing — derived from Equation 10 using coefficients from Figure 6. Estimates are shown for two post-treatment horizons: the medium run (0–30 years) and the long run (40–60 years). MWTP values represent the percentage increase in monthly rent a household would be willing to pay for a one-unit change in the respective attribute (Columns 2 and 4). Changes are normalized as follows: 100 units for Apartments, $1,000 for Construction Costs, one storey for Height, and 10 percentage points for Ground Coverage and Racial Composition. Rent values are based on the average median contract rent (in 2010 dollars) by ring and time period (Columns 3 and 5). Each estimate is reported separately for public housing tracts (PH tracts) and adjacent areas (1st ring). The sample includes only projects constructed before 1970. Standard errors for the MWTP estimates are computed from the empirical variance of the underlying population and rent coefficients, as detailed in Appendix A1.
Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

the Park model. Tall public housing buildings are consistently viewed as undesirable by private households—particularly among Black and White residents. This suggests that vertical scale carries strong negative connotations, potentially related to safety concerns or social stigma. By contrast, the open-space layout created by low ground coverage appears to have little impact on neighborhood preferences among private households.

I next examine how private households respond to changes in the racial composition of public housing residents—measured as the change from the initial tenant composition at project opening to 2010. The results show strong evidence of homophilic preferences, particularly among White residents. A 10 percentage point increase in the share of White public housing residents is associated with a statistically significant increase in MWTP among White households not living in public housing: \$276.71 in the medium run and \$254.78 in the long run for tracts with public housing. Positive MWTP is also observed in adjacent tracts — \$75.73 in the medium run and \$164.55 in the long run — suggesting that preferences for same-race neighbors extend beyond directly treated neighborhoods.

By contrast, Black households exhibit a negative MWTP for living near public housing with a higher share of Black tenants. A 10 percentage point increase in the share of Black public housing residents is associated with a statistically significant reduction in MWTP of –\$369.63, indicating that Black residents would require this amount in monthly rent compensation to live near such developments. At the same time, Black households show a strong positive MWTP for living near Hispanic public housing residents, with a long-run estimate of \$213.26. Interestingly, Black MWTP for adjacent tracts experiencing increases in White public housing residents is initially negative in the medium run (–\$145.74), though this effect fades and becomes statistically insignificant over time.

For Hispanic households, MWTP estimates are generally imprecise. However, one notable pattern emerges: in the long run, a 10 percentage point increase in the share of Black public housing residents is associated with a statistically significant MWTP decline of –\$116.41, indicating some aversion to areas with increasing concentrations of Black public housing tenants.

These results broadly align with the MWTP estimates for later-constructed projects, reported in Appendix D5 (see Table D5.1). Across both early and later cohorts, households consistently place positive value on higher construction costs, reinforcing the interpretation that private residents prefer better-built public housing. Conversely, larger-scale developments are generally viewed unfavorably, with negative MWTP estimates across all groups, suggesting concerns related to density, perceived stigma, or lower neighborhood quality.

The patterns also support three broader conclusions. First, private households systematically favor higher-quality public housing, as captured by construction costs. Second, there is evidence of racial homophily and selective neighborhood preferences: White households exhibit strong positive MWTP for areas with increasing shares of White public housing residents, while Black households prefer areas with more Hispanic tenants and avoid areas with growing concentrations of Black tenants. hese patterns reflect both co-ethnic preferences and aversion to stigma. Third, taller buildings are broadly disfavored by all groups, lending support to longstanding critiques of the *Tower in the Park* model. While the precise interpretation of building height remains open—potentially reflecting perceptions of crime, isolation, or socioeconomic stigma—its consistently negative valuation points to a clear preference among private residents for lower-rise public housing designs.

9 Conclusion

This paper provides new evidence on the long-run effects of public housing construction on neighborhood composition and housing markets in New York City. Leveraging archival data and distinguishing between public housing tenants and privately housed residents, I document persistent demographic and price changes unfolding over decades. Projects built before 1970 led to substantial outflows of White and Black private residents, and large rent declines—evidence of a strong market response. In contrast, post-1970 projects had limited effects, primarily reflecting the direct placement of residents, rather than inducing broader neighborhood adjustment.

To move beyond average effects, I embed the reduced-form estimates in a discrete location choice model, and recover marginal willingness to pay (MWTP) for key public housing attributes. The results show that higher construction costs—used as a proxy for quality—are positively valued by all groups. In contrast, high-rise *Tower in the Park* designs consistently reduce neighborhood desirability, especially among White and Black residents. Racial composition also matters. White households strongly prefer areas with more White public housing tenants, while Black households exhibit more heterogeneous, and sometimes negative, preferences for co-ethnic clustering.

These findings highlight a core policy trade-off. While large-scale developments expand the housing stock, they risk triggering neighborhood flight and rent declines. In contrast, smaller, higher-quality projects are more likely to enhance neighborhood appeal but come at higher public costs. Designing effective affordable housing policy thus requires not only more housing, but better housing—tailored to how physical design and resident composition shape long-run neighborhood dynamics.

Future research should explore optimal public housing scale and placement by modeling resident preferences and planner trade-offs. Additional work linking building attributes to crime or renovation needs could help disentangle whether negative responses reflect stigma or real externalities.

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Online Appendix for "Public Housing Preferences: Evidence from New York City 1930-2010"

Maximilian Guennewig-Moenert

Appendix Contents

A Model Details

This section displays the full model from Section 7. The city consists of ethnic groups $g \in \{B, W, H\}$ where $\sum N^g = N$ is the total number of residents in New York City. Individual i's utility depends on consumption of a single city-wide final good, housing units, residential amenities, and an individual-specific idiosyncratic shock that varies with residence location. Under Cobb Douglas preferences, the utility of individual i of group g living in tract k is given by:

$$U^g_{ik} = f\left(B^g_{ik}, \epsilon^g_{ki}\right) \cdot \left(\frac{C_i}{\alpha}\right)^{\alpha} \left(\frac{h_{ik}}{1-\alpha}\right)^{1-\alpha} \tag{A.1}$$

Group-specific residential amenities B_{gk} capture common features that make a location a more or less desirable place to live. The consumption good C_i is chosen as numeraire. I parameterize $f\left(B^g_{ik}, \epsilon^g_{ki}\right)$ with an exponential function and assume that public housing affects amenities by distance:

$$f\left(B^g_{ik}, \epsilon^g_{ki}\right) = \exp\left(\beta^g_{PH_1} PH_{1,k} + \beta^g_{PH_2} PH_{2,k} + \epsilon^g_{ik}\right) \tag{A.2}$$

The $\beta^g_{PH_1}$ and $\beta^g_{PH_2}$ are preference parameters for indicator variables, denoting whether a tract k contained a public housing project, $PH_{1,k} = \mathbb{1}[k \in PH]$, or k was a neighbor of a tract $j \neq k$ with a public housing project, $PH_{2,k} = \mathbb{1}[k \in \text{1st ring}]$. Log indirect utility is then given by:

$$V^g_{ik} = \varphi^g_{ik} + \epsilon_{ik} \tag{A.3}$$

where φ^g_{ik} is the component of indirect utility for census tract k that is common to all individuals of group g — called mean indirect utility hereafter, and ϵ_{ik} is an idiosyncratic shock drawn from an extreme value type I distribution. The common component of indirect utility is:

$$\varphi^g_{ik} = \beta^g_{PH_1} PH_{1,k} + \beta^g_{PH_2} PH_{2,k} + \log(w^g_k) - \alpha \log(r_k) \tag{A.4}$$

Given the distributional assumption on ϵ_{kt}, the probability that an individual i of group g chooses to live in tract k is:

$$\pi^g_{ik} = \frac{\exp(\varphi^g_{ik})}{\sum_{j=1}^{M} \exp(\varphi^g_{ij})} \tag{A.5}$$

It should be noted that the denominator is type-specific but not location-specific. It measures the expected utility of living in the city and is treated as a constant. To see this, define $\bar{v}^g = \frac{1}{C}\sum_{k=1}^{C}\exp\left(\frac{V_k^H}{\sigma^H}\right)$, where C is the number of choices, and let $a^g = C\bar{v}^g$. For a given C that is large enough, any change in φ_{ik}^g does not affect $\bar{v}^g$.

The demand for living in neighborhood k equals the total number of individuals, across all groups, that want to live in k. The total population of group g, N^g, in New York City is treated as exogenous, yielding the following housing demand equation:

$$D_k = \sum_g \pi_k^g N^g \tag{A.6}$$

The model is closed by assuming an isoelastic housing supply function such that the number of housing units in tract k is given by:

$$S_k = \delta_k r_k^{\phi_k} \tag{A.7}$$

where δ_k is a supply shifter and ϕ_k is the tract-specific supply elasticity. Imposing the market-clearing condition $S_k = D_k$ closes the model and allows us to solve for equilibrium rents. Equilibrium rent prices are then given by:

$$\log(r_k^*) = \frac{1}{\phi_k + \alpha}\left[\log\left(\sum_g \frac{\exp\left(\beta_{PH_1}^g PH_{1,k} + \beta_{PH_2}^g PH_{2,k} + \log(w_k^g)\right)}{\bar{v}^g} N^g\right) - \log(\delta_k)\right] \tag{A.8}$$

Note that I assume the expenditure share of housing does not vary across groups. Using Equation A.8, one can solve for equilibrium population of blacks and whites in tract k:

$$\log(N_k^{g*}) = \beta_{PH_1}^g PH_{1,k} + \beta_{PH_2}^g PH_{2,k} + \log(w_k^g) - \alpha\log(r_k^*) - \log(\bar{v}^g) + \log(N^g) \tag{A.9}$$

I derive two comparative statics from Equation A.8 and Equation A.9 to evaluate the equilibrium response of rent prices and population to public housing. First, I differentiate equilibrium rents with respect to PH_R, where $R \in \{PH\ tract, \text{1st ring}\}$ indicates the distance relationship to public housing:

$$
\begin{aligned}
\frac{d\log(r_k^*)}{dPH_{R,k}} &= \frac{1}{\phi_k+\alpha}\cdot\frac{1}{\sum\limits_g \frac{\exp\left(\beta^g_{PH_1}PH_{1,k}+\beta^g_{PH_2}PH_{2,k}+\log(w_k^g)\right)}{\bar{v}^g}N^g} \\
&\quad\times\sum_g \frac{\beta^g_{PH_R}\exp\left(\beta^g_{PH_1}PH_{1,k}+\beta^g_{PH_2}PH_{2,k}+\log(w_k^g)\right)}{\bar{v}^g}N^g \\
&= \frac{1}{\phi_k+\alpha}\cdot\frac{\sum\limits_g \beta^g_{PH_R}\tilde{T}_g}{\sum\limits_g \tilde{T}_g} \\
&= \frac{\Xi^g_R}{\phi_k+\alpha}
\end{aligned} \tag{A.10}
$$

where $\tilde{T}_g = \frac{\exp\left(\beta^g_{PH_1}PH_{1,k}+\beta^g_{PH_2}PH_{2,k}+\log(w_k^g)\right)}{\bar{v}^g}N^g$. Equation A.10 reveals that rent prices are determined by the income share spent on housing, the local housing supply elasticity, and probability-weighted preference parameters.

Using Equation A.10 gives the equilibrium population response of group g in tract k to the construction of public housing:

$$
\frac{d\log(N_k^{g*})}{dPH_{R,k}} = \beta^g_{PH_R} - \alpha\cdot\frac{\partial\log(r_k^*)}{\partial PH_{R,k}} = \beta^g_{PH_R} - \alpha\cdot\frac{\Xi^g_R}{\phi_k+\alpha} \tag{A.11}
$$

Equation A.10 and Equation A.11 give an expression for the change in equilibrium rents and tract population in treated and adjacent tracts.

A1 Variance of the Marginal Willingness to Pay

Next, I show how to obtain variance estimates for Marginal Willingness to Pay (MWTP) for a given public housing attribute. First, note that from Equation A.11 the preference parameter can be identified using the reduced form coefficients from the empirical model:

$$
\begin{aligned}
\beta^g_{PH_R,t} &= \frac{\partial\log N_{kt}^{g*}}{\partial PH_{Rkt}} - \alpha\times\frac{\partial\log r_{kt}^*}{\partial PH_{Rkt}} \\
&= \hat{\beta}^g_{PH_R,t} + \alpha\times\hat{\beta}^{rent}_{PH_R,t}
\end{aligned} \tag{A.12}
$$

where $\hat{\beta}^g_{PH_r,t}$ corresponds to the causal effect of public housing on ethnic group $g\in\{W,B\}$ and $\hat{\beta}^{rent}_{PH}$ to the effect on rental prices. Given that in the empirical model I am interested in the time-varying effects of public housing, I index the model parameters by $t\in\{0-30\text{ years}, 40-60\text{ years}\}$ to denote the medium and long run. Equation A.12 shows that the structural preference parameter can be recovered from reduced-form coefficients on population and rent using the equilibrium condition. The variance of $\beta^g_{PH_R,t}$ is then given by:

$$\text{Var}(\beta^g_{PH_R,t}) = \text{Var}(\hat{\beta}^g_{PH_R,t}) + \alpha^2 \text{Var}(\hat{\beta}^{rent}_{PH_R,t}) \tag{A.13}$$

Note that since $\hat{\beta}^g_{PH_R,t}$ and $\hat{\beta}^{rent}_{PH_R,t}$ are estimated from two independent regressions. I assume that $\text{COV}(\hat{\beta}^g_{PH_R,t}, \hat{\beta}^{rent}_{PH_R,t}) = 0$. To link these preferences to economic valuation, I compute the marginal willingness to pay (MWTP) implied by the structural preference parameters. To assess the MWTP for specific public housing attributes, I interact the public housing dummy, $PH_{r,k}$, with a vector of public housing attributes **ATTR**. Using indirect utility from Equation A.3, the MWTP for a location $r \in \{PH\ tract, \text{1st ring}\}$ due to a change in a public housing attribute is defined as:

$$MWTP^g_{ATTR_{i,t}}\Big|_{PH_R=R} = -\frac{dV/dATTR_{i,t}}{dV/d\log r_{k,t}} = \frac{\beta^g_{ATTR_{i,t}}}{\alpha} \tag{A.14}$$

where $\beta^g_{ATTR_{i,t}}$ is the average preference parameter of a household from group g for public housing attribute i. Since the derivative in Equation A.14 is with respect to $\log r_{k,t}$, MWTP represents the percentage increase in rent for a marginal change in $ATTR_i$. Thus, the dollar rent change is $dr_{k,t}/dATTR_{i,t} = r_{k,t} \times MWTP^g_{ATTR_{i,t}}$. It follows that the variance of $MWTP^g_{ATTR_{i,t}}$ is given by:

$$\text{Var}\left(MWTP^g_{ATTR_{i,t}}\right) = \frac{1}{\alpha^2}\left[\text{Var}(\beta^g_{ATTR_{i,t}})\right] \tag{A.15}$$

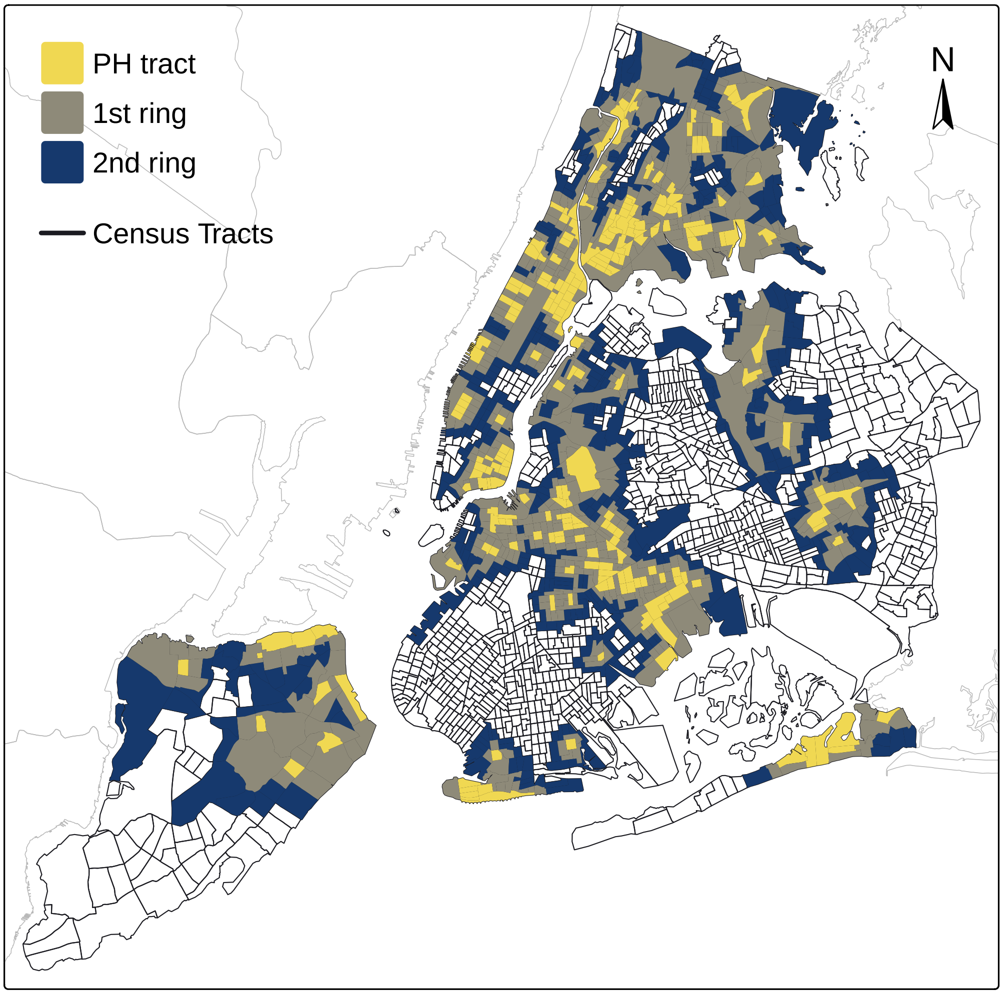
To assess statistical significance, I assume that the MWTP estimates follow an approximate normal distribution, based on the delta method. This standard approach allows estimation of standard errors for nonlinear transformations of estimated parameters. I report two-sided p-values based on the standard normal distribution, as shown in Equation A.16:

$$p = 2 \cdot \Phi\left(-\left|\frac{\widehat{\text{MWTP}}}{\text{SE}(\widehat{\text{MWTP}})}\right|\right) \tag{A.16}$$

where Φ is the cumulative distribution function of the standard normal distribution,

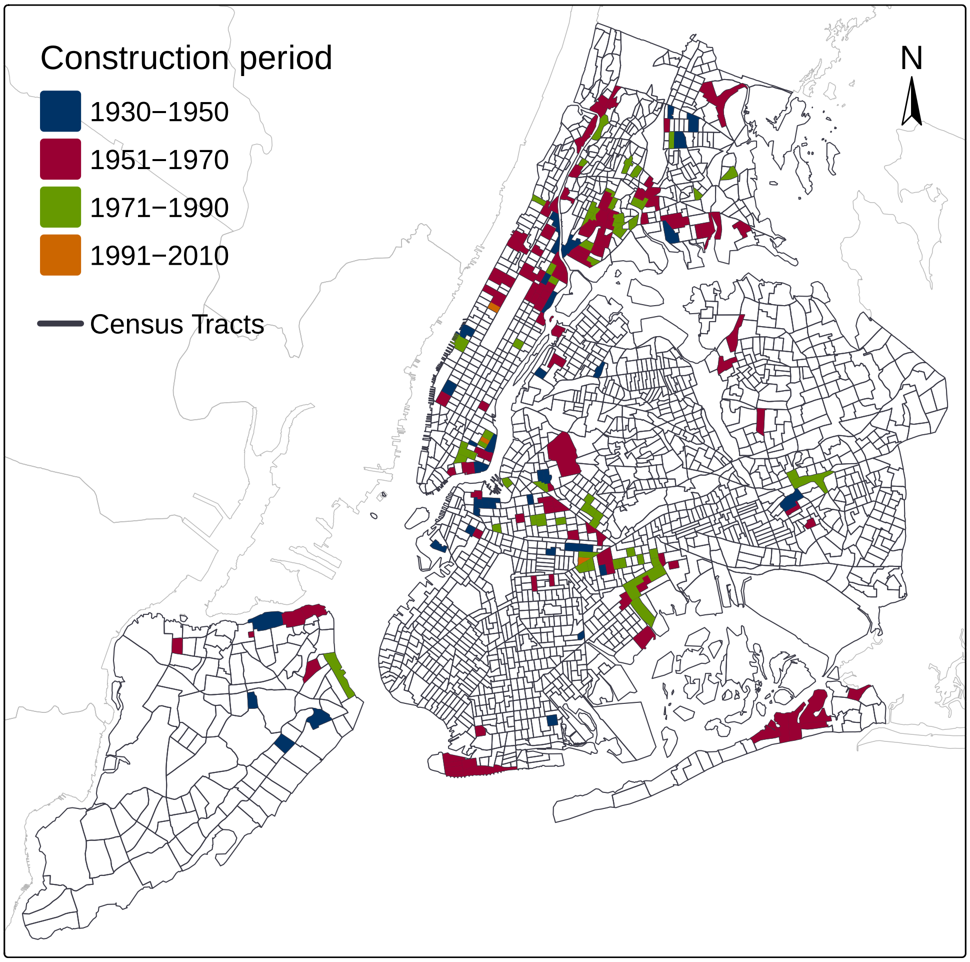
B Maps

Figure B.1: Tracts by Distance Relationship to Public Housing



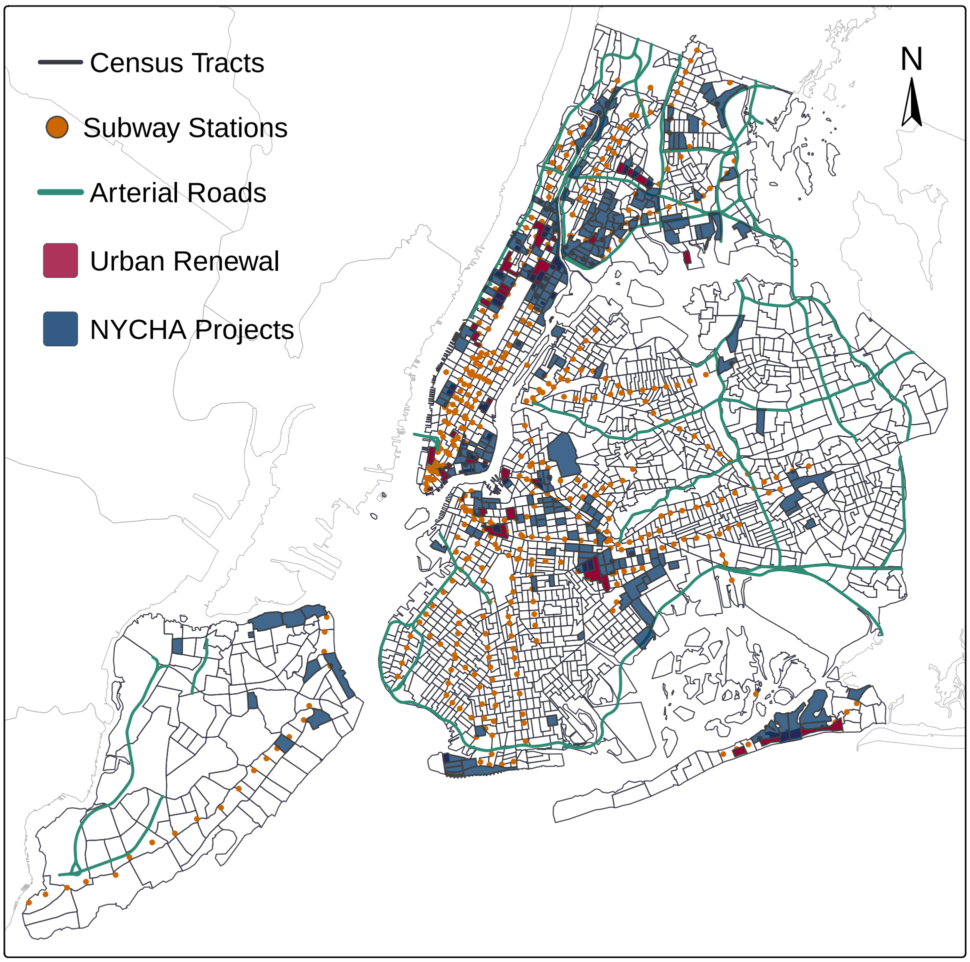
Note. Tracts by distance relationship as used in the analysis in panel setup.

Figure B.2: Evolution of Public Housing by Construction Period



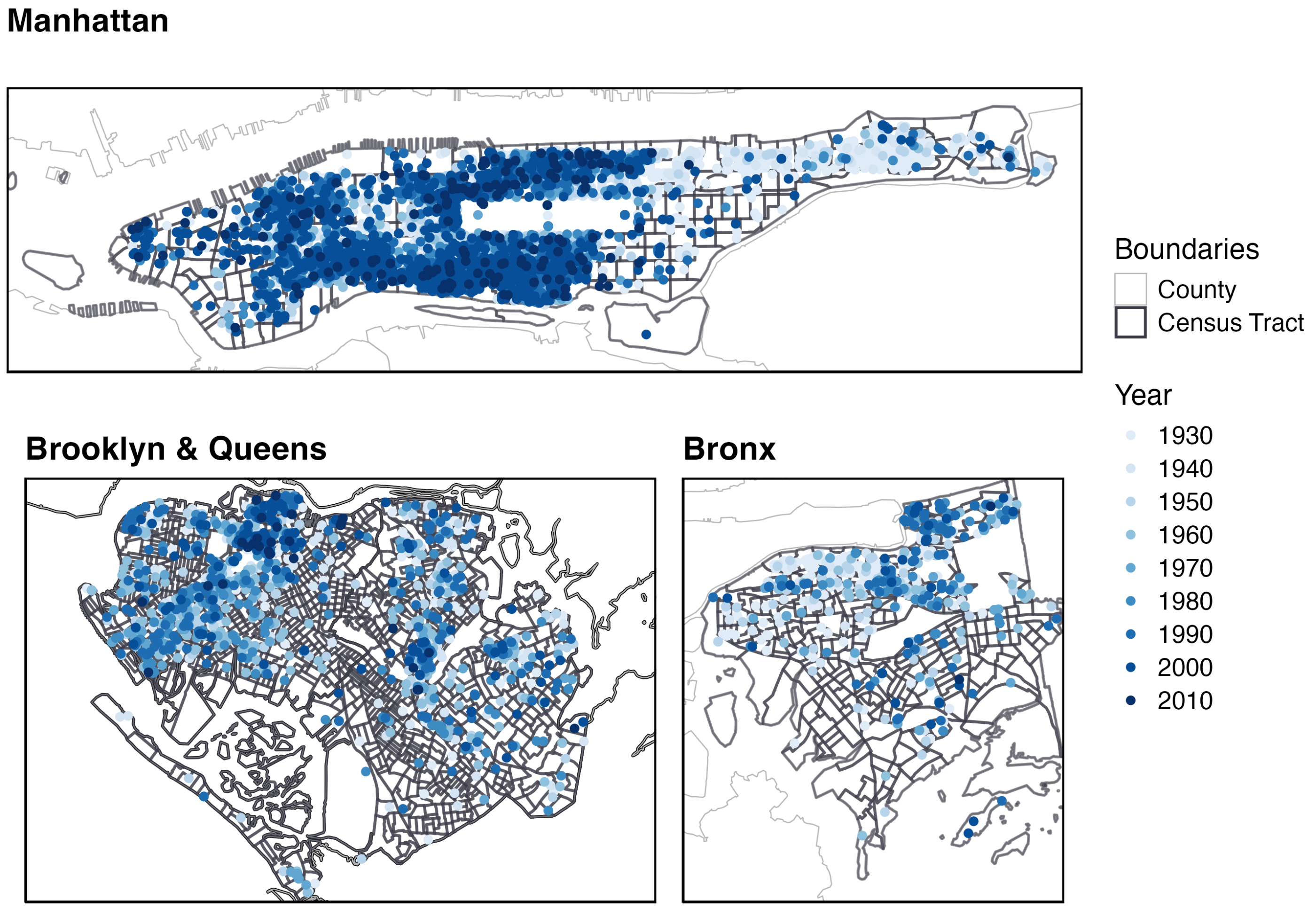
Note: Figure B.2 displays 2010 census tract boundaries in New York City. Tracts highlighted in color contain at least one public housing project. Colors indicate the construction period of the first completed project in each tract, grouped into four categories: 1930–1950 (blue), 1951–1970 (red), 1971–1990 (green), and 1991–2010 (orange). Some tracts include multiple projects constructed at different times, but classification is based on the earliest project. The map illustrates the spatial and temporal spread of public housing development across the city.
Source. La Guardia and Wagner Archives, NYCHA development data book. Details on the construction of the data set can be found in Section 3.

Figure B.3: Public Housing Locations and Urban Infrastructure in New York City



Note: Figure B.3 displays the geographic distribution of NYCHA public housing projects across New York City. The map overlays key infrastructure and planning features, including census tract boundaries, subway stations (orange), arterial roads (emerald), and designated urban renewal areas (red).

Figure B.4: Spatial Distribution of Rental Properties



Note: Figure B.4 shows the spatial distribution of geocoded rental listings from the *New York Times* real estate section across four boroughs: Manhattan, Brooklyn, Queens, and the Bronx. Each point represents a rental property, shaded by decade from 1930 to 2010. Boundaries correspond to census tracts (dark grey) and county (light grey) lines.
Source. NYT Real Estate Section. Details on the construction of the data set can be found in Section 3 and Appendix C4.

C Data

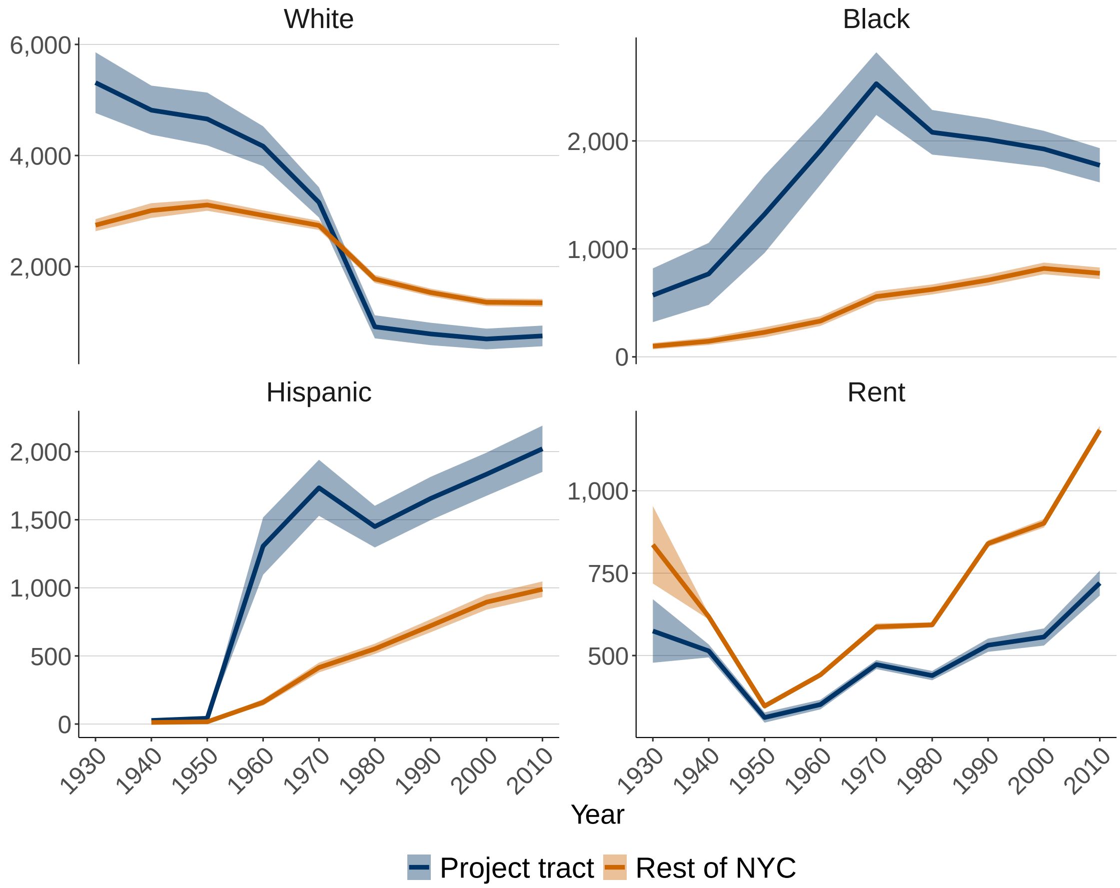
Table C.1: Definition of Variables and Data Sources

	Years	Level	Description	Source
Asking rents	1930–2010	Property	Asking rent and property characteristics.	New York Times Real Estate Section; `proquest.com`
Med. contract rent	1930–2010	Tract	Median contract rent (2010 $) at the census-tract level.	Manson et al. (2024)
White population	1930–2010	Tract	White population at the census-tract level.	Manson et al. (2024)
Black population	1930–2010	Tract	Black population at the census-tract level.	Manson et al. (2024)
Black public-housing residents	1930–2010	Project	Black residents at project opening date.	NYCHA Resident Data Book; LaGuardia & Wagner Archives
White public-housing residents	1930–2010	Project	White residents at project opening date.	NYCHA Resident Data Book; LaGuardia & Wagner Archives
Project height	1930–2010	Project	Number of storeys of each public-housing project.	NYCHA Development Data Book
Ground coverage	1930–2010	Project	Ground-floor area / project area (%).	NYCHA Development Data Book
Number of public-housing units	1930–2010	Project	Number of apartments in each public-housing project.	NYCHA Development Data Book
Arterial street	1930–2010	Segment	NYC arterial streets with opening dates.	NYC Open Data; `nycroads.com`
Urban-renewal programmes	1930–2010	Area	Areas designated for urban renewal and start date.	Digital Scholarship Lab (2018)
Subway stations	1930–2010	Station	NYC subway stations with opening dates.	NYC Open Data; Wikipedia

Note. Table C.1 lists all variables used in the analysis, along with their temporal coverage, geographic level, definitions, and data sources. Information on public housing residents at the time of project opening is drawn from the LaGuardia & Wagner Archives for projects constructed up to 1970, and from the NYCHA Resident Data Book for projects completed afterward. Subway and arterial road shapefiles are obtained from NYC Open Data, while opening dates are sourced from Wikipedia and `nycroads.com`.

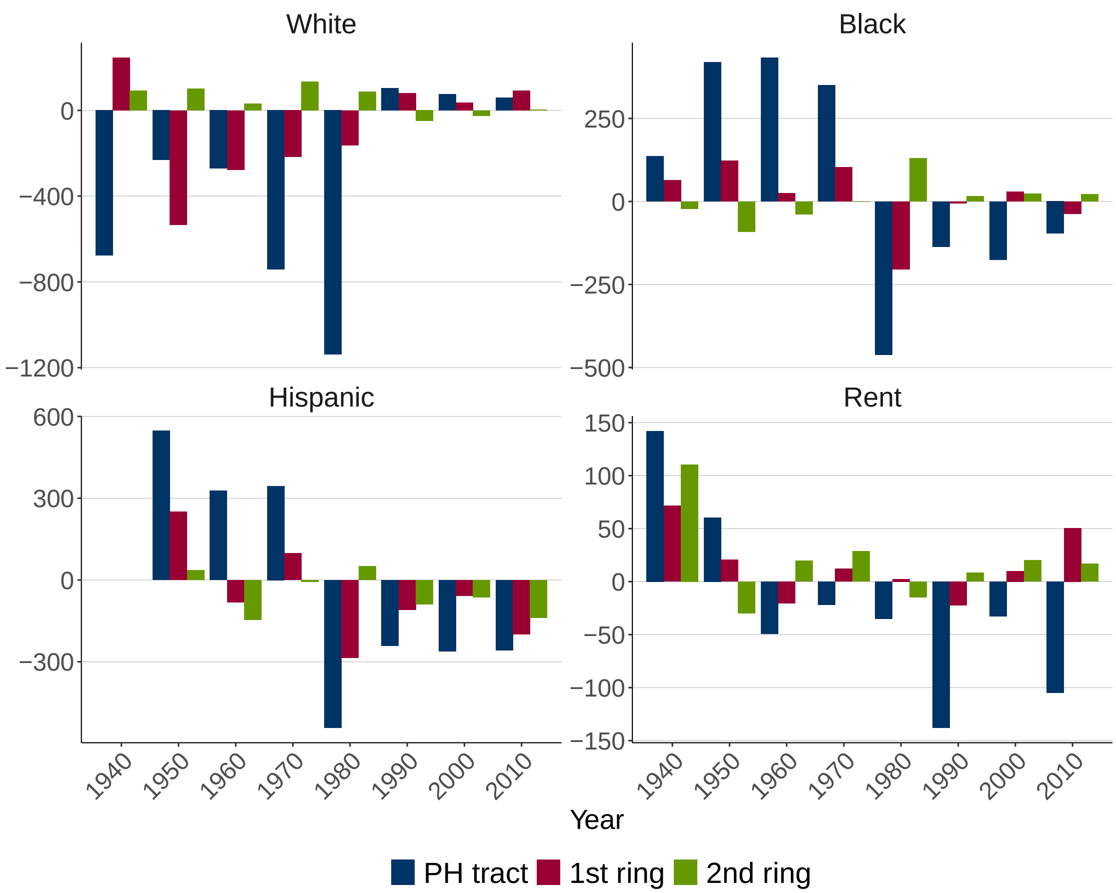
C1 Census Statistics

Figure C1.1: Demographic Trends



Note. Figure C1.1 reports trends of the main outcome variables. It shows yearly averages for: White and Black population, and median contract rent in New York City. I compute averages for all treated tracts, or, in other words, those which ever had a public housing unit within its boundaries (Project Tract) and all remaining tracts in New York City (Rest of NYC).

Figure C1.2: Deviation from Average City Tract



Note. Figure C1.2 reports the deviation of the average a treated (PH tract), adjacent (1st ring) and control tract (2nd ring) as defined in Section 4 from the average tract in the rest of new york city.
Source. US Decennial Census. Details on construction of the data set can be found in section 3.

C2 Public Housing Statistics

Table C2.1: Public Housing Characteristics by Construction Decade

	1940	1950	1960	1970	1980	1990	2000
	Total counts						
Projects	9	28	58	80	63	26	9
Units	10244	30382	70601	41933	14911	5715	626
	Median characteristics						
Stories	6	10	14	18	12	7	6
Units	1531	1156	1246	441	221	189	51
Ground coverage	29.74%	18.86%	14.50%	17.62%	33.47%	28.01%	42.68%
Construction cost	$12'854.93	$14'712.01	$13'963.55	$15'409.28	$17'287.91	$25'965.57	$26'962.55
	Average characteristics						
Stories	5.33	9.43	12.78	16.95	12.27	8.12	5.44
Units	1138.22	1085.07	1217.26	524.16	236.68	219.81	69.56
Ground coverage	27.87%	18.29%	14.92%	21.61%	31.86%	28.26%	46.71%
Construction cost	$18'622.04	$16'744.14	$16'626.30	$17'682.21	$17'123.95	$25'097.79	$29'298.38

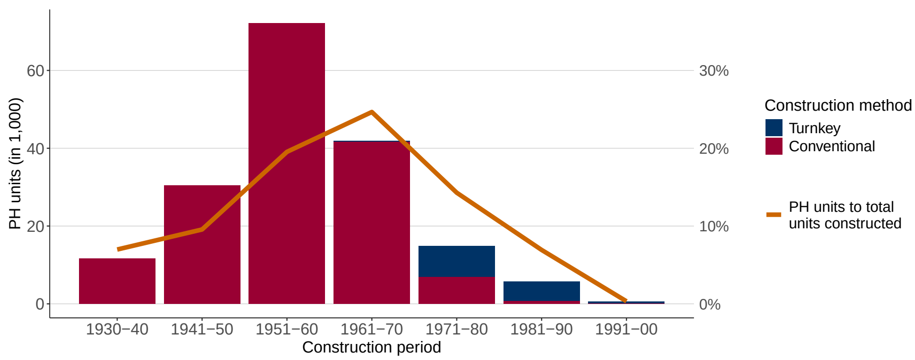
Note. Table C2.1 displays public housing project information within the decade of their construction; projects were grouped into construction period cohorts based on their opening date. The first two rows report total counts by construction decade. Row three to six shows median public housing characteristics by construction decade. Rows seven to ten average characteristics. The average was taken across all public housing projects constructed within a decade. Construction cost refer to construction cost per room, are deflated by the CPI and given in $2010.
Source. NYCHA Development Data Book. Details on construction of the data set can be found in Section 3.

Table C2.2: Correlation Matrix of Public Housing Attributes

	# Units	Init. % White	Init. % Black	Init. % Hispanics	Const. Cost	# Stories	% Coverage
# Units		0.40***	-0.24***	-0.12*	-0.12*	0.24***	-0.57***
Init. % White	0.40***		-0.75***	-0.52***	0.06	0.07	-0.41***
Init. % Black	-0.24***	-0.75***		0	-0.16**	-0.06	0.23***
Init. % Hispanics	-0.12*	-0.52***	0		0.03	0.16**	0.06
Const. Cost	-0.12*	0.06	-0.16**	0.03		-0.03	0.16**
# Stories	0.24***	0.07	-0.06	0.16**	-0.03		-0.48***
% Coverage	-0.57***	-0.41***	0.23***	0.06	0.16**	-0.48***	

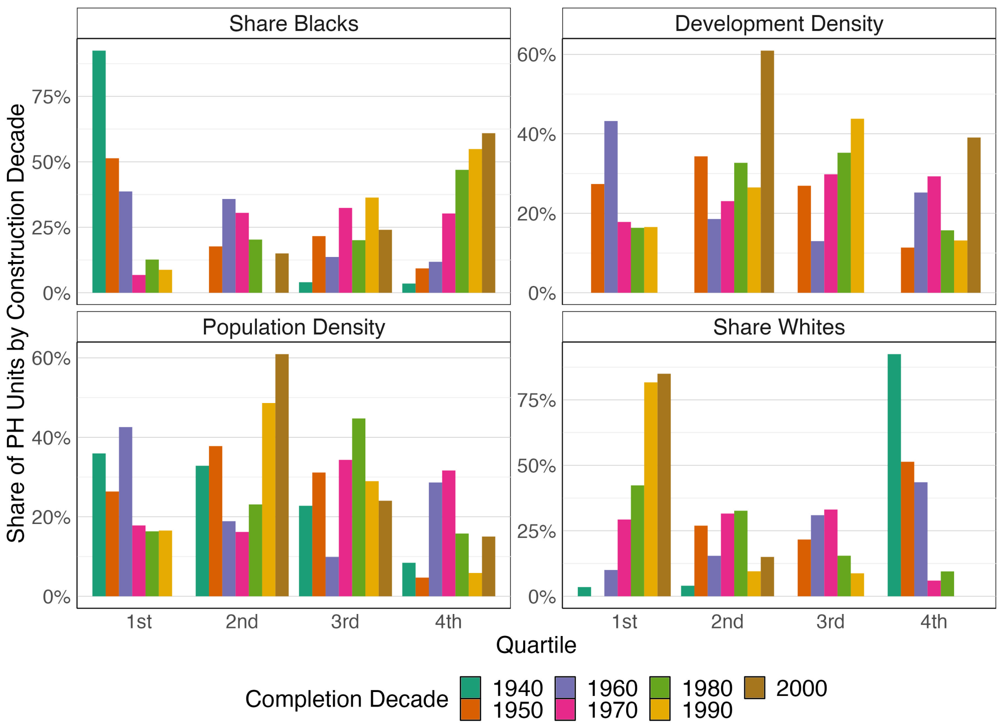
Note. Table C2.2 displays the pairwise correlation matrix of attributes used in the main analysis for all newly constructed public housing projects. The table includes measures such as building height, number of units, construction costs, and initial racial composition.
Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Figure C2.1: Evolution of Public Housing by Construction Period



Note. Figure C2.1 reports trends of public housing by construction decade. Projects have been grouped in construction periods by their completion date. The figure shows the total number of units within a decade. There are two acquisition methods. Under the *Conventional Method*, the authority acquires the land and contracts for General Construction, Heating and Ventilation, Elevators, Electrical, and Plumbing work. Under the *Turnkey Method*, the developer buys the land, constructs the development, and sells it to the Authority under the terms of a pre-agreed contract. The orange line shows the total number of public units as a share of total units constructed in New York City within the decade.
Source. NYCHA development data book. Details on the construction of data the data set can be found in Section 3.

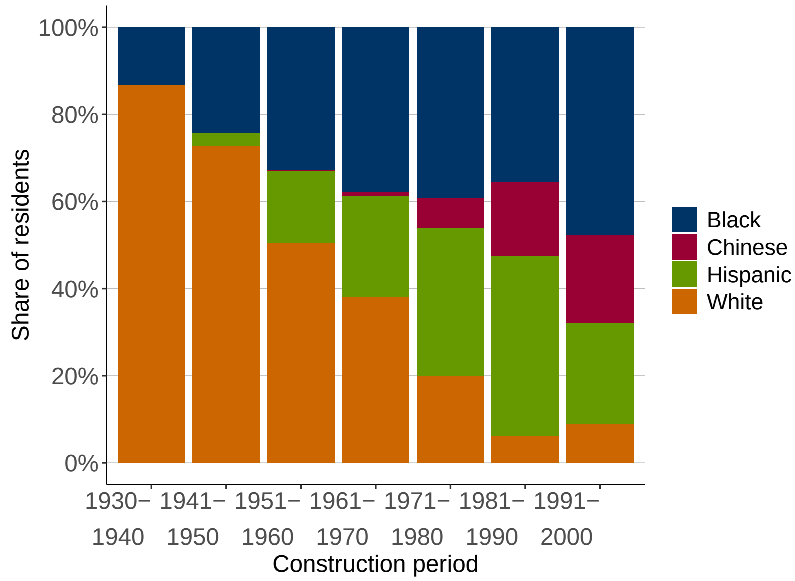
Figure C2.2: Public Housing Units by Construction Cohort and Quartile of Baseline Tract Characteristics



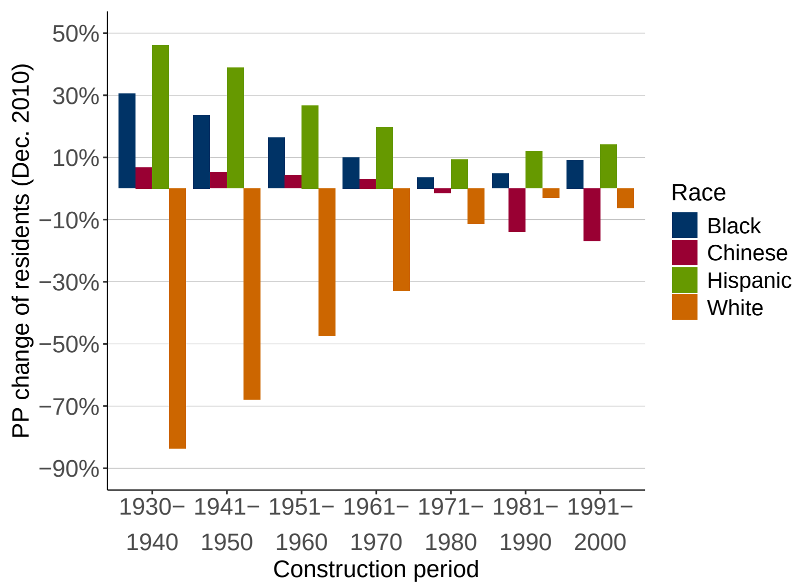
Note: Figure C2.2 shows the share of public housing units by construction cohort by quartile of baseline tract characteristics. Tract characteristics were taken the decade before a public housing project arrived. Next, total public housing was grouped by quartile as a share of the total number of units constructed within the decade. Each decade refers to the projects constructed nine years before. Details on Data construction *Source.* NYCHA development data book and US federal census. Details on the construction of the data set can be found in Section 3.

Figure C2.3: Racial Composition by Construction Decade

(a) Initial allocation of residents



(b) Change of residents



Note. Figure C2.3 displays the ethnic composition of NYCHA projects based on their construction decade. Projects have been grouped in construction periods by their completion date. Panel a presents the resident shares by ethnicity at the time of initial occupancy. Panel b illustrates change for each ethnic group from the project's start date to December 2010 in percentage points.
Source. La Guardia and Wagner Archives, NYCHA development data book. Details on the construction of the data set can be found in Section 3.

Figure C2.4: Archival Source for Project-Level Racial Data

RACIAL DISTRIBUTION IN OPERATING PROJECTS

AT INITIAL OCCUPANCY AND ON DECEMBER 31, 1956

PROJECT	Completion of Initial Occupancy Month/Year	AT INITIAL OCCUPANCY									
		White		Negro		Chinese		Other *		Total	
		No.	%	No.	%	No.	%	No.	%	No.	%
CITY PROGRAM - PART I											
Elliott	July, 1947	555	92.7	36	6.0	-	-	8	1.3	599	100.0
First	May 1936	122	100.0	-	-	-	-	-	-	122	100.0
Riis City	Jan. 1949	528	91.3	35	6.1	4	0.7	11	1.9	578	100.0
Vladeck City	Oct. 1940	236	98.3	3	1.3	-	-	1	0.4	240	100.0
TOTAL		1,441	93.6	74	4.8	4	0.3	20	1.3	1,539	100.0
CITY PROGRAM - PART II											
Colonial Park	Oct. 1951	23	2.4	876	90.3	1	0.1	70	7.2	970	100.0
Eastchester	June 1950	761	87.6	95	11.0	-	-	13	1.4	869	100.0
Sheepshead Bay	Aug 1950	1,018	96.9	32	3.0	-	-	1	0.1	1,051	100.0
South Beach	March 1950	385	92.1	25	5.9	1	0.3	7	1.7	418	100.0
Woodside	Dec. 1949	1,243	91.9	103	7.6	1	0.1	6	0.4	1,353	100.0
TOTAL		3,430	73.6	1,131	24.2	3	0.1	97	2.1	4,661	100.0
CITY PROGRAM - PART III											
Arverne	Feb. 1951	390	94.4	23	5.6	-	-	-	-	413	100.0
Berry	Oct. 1950	457	90.7	45	8.9	1	0.2	1	0.2	504	100.0
Boulevard	Mar. 1951	1,288	89.8	143	10.0	1	0.1	1	0.1	1,433	100.0
Dyckman	April 1951	933	80.5	211	18.2	2	0.2	13	1.1	1,159	100.0
Glenwood	July 1950	1,139	96.1	45	3.8	-	-	1	0.1	1,185	100.0
Gun Hill	Nov. 1950	622	85.3	101	13.9	-	-	6	0.8	729	100.0
Lexington	Mar. 1951	148	33.5	278	62.9	-	-	16	3.6	442	100.0
Marble Hill	Mar. 1952	1,289	76.8	350	20.9	1	0.1	37	2.2	1,677	100.0
Nostrand	Dec. 1950	1,104	96.2	39	3.4	1	0.1	3	0.3	1,147	100.0
Parkside	June 1951	806	92.1	63	7.2	1	0.1	5	0.6	875	100.0
Pelham	July 1950	1,200	95.3	54	4.3	-	-	5	0.4	1,259	100.0
Pomonok	June 1952	1,813	87.9	238	11.5	-	-	13	0.6	2,064	100.0
Ravenswood	July 1951	1,902	88.0	252	11.7	1	0.0	6	0.3	2,161	100.0
Sedgwick	Mar. 1951	671	85.7	105	13.4	1	0.1	6	0.8	783	100.0
Todt Hill	June 1950	455	91.2	42	8.4	1	0.2	1	0.2	499	100.0
TOTAL		14,217	87.0	1,989	12.2	10	0.1	114	0.7	16,330	100.0
CITY PROGRAM - PART IV											
Bay View	June 1956	1,492	92.8	102	6.4	-	-	13	0.8	1,607	100.0

Note. Figure C2.4 shows page three of the Dec. 31st, 1956 NYCHA report "Racial Distribution in Operating Projects", indicating the racial distribution at initial occupancy for each project.
Source. LaGuardia and Wagner Archives, NYCHA collection, Box. Nr. 0071B6.

C3 Tract Harmonisation

A major challenge in using census tract-level data for longitudinal analysis is that tract boundaries change substantially over time, making it difficult to construct a time-consistent panel dataset. I address this issue by reweighting observations based on overlapping area weights, using 2010 census tract boundaries as the target geography. Specifically, I aggregate count variables such as white and black population—using this method. For 1930 I use the full population count census census fom IPUMS which I aggregate in enmueration (ED) district level. Thus, instead of census tract, I use overlapping area weights for enumeration districts. Let i index 2010 census tracts and j index historical tracts from earlier census years. For each 2010 tract i, the weighted estimate $\hat{y}_i$ for a given outcome variable y is calculated as:

$$\hat{y}_i = \sum_{j \in \mathcal{J}_i} \frac{A_{ij}}{A_j} \cdot y_j,$$

where y_j is the value of the outcome variable in historical tract j, A_{ij} is the area of intersection between historical tract j and 2010 tract i, and A_j is the total area of tract j. The set $\mathcal{J}_i$ contains all historical tracts j that overlap with tract i.

This method assumes that the outcome variable is uniformly distributed within each historical tract, which may introduce error. For instance, if tract j in year t contains a spatially concentrated low-income population, but overlaps evenly with multiple 2010 tracts, the method would misallocate residents uniformly across those tracts—biasing spatial estimates.

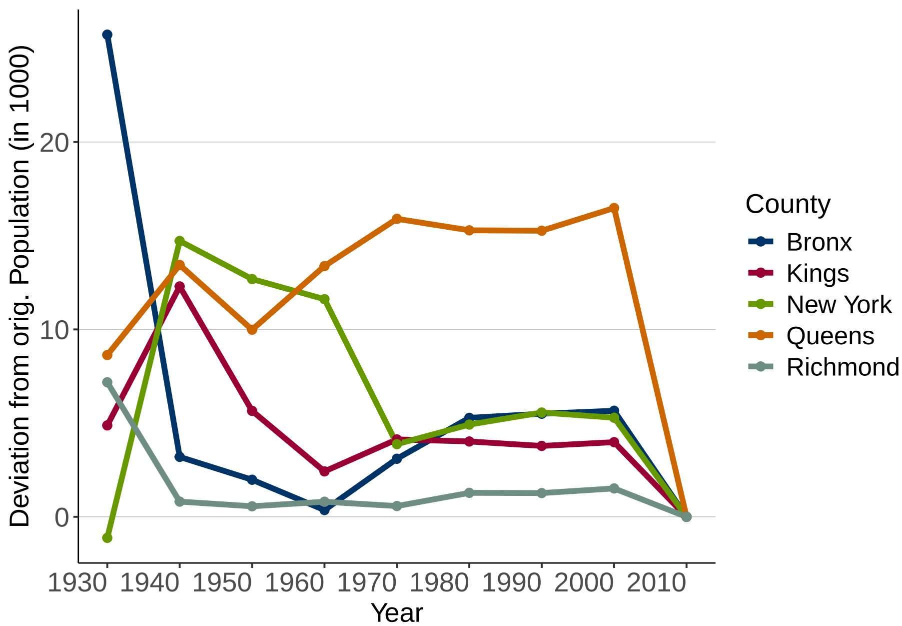
For variables such as median rent and household income, which are not count-based, I calculate tract-level estimates as weighted averages across historical geographies using overlapping area weights. Specifically, for each 2010 census tract i, I compute:

$$\hat{y}_i = \frac{\sum_{j \in \mathcal{J}_i} y_j \cdot w_{ij}}{\sum_{j \in \mathcal{J}_i} w_{ij}},$$

where w_{ij} is the proportion of area of historical tract j overlapping with 2010 tract i, conditional on $w_{ij} > 0.05$. This approach accommodates spatial uncertainty while preserving information from all overlapping historical tracts. A fallback value from the highest-overlap tract is also computed, but not used in the main analysis.

Figure C3.1 compares the reweighted population series at the borough level to the original series. While overall trends are preserved, deviations from the original data vary across boroughs and over time. The largest deviations occur particularly for Queens, where cumulative differences exceed 10,000 residents in some years. This reflects more complex boundary changes—such as tract splits and consolidations—that introduce greater error

Figure C3.1: Deviation due to Boundary Harmonization



Note. Figure C1.2 shows the difference between total population aggregated from original census tracts and total population estimated using area-weighted aggregation based on overlapping boundaries, by borough. *Source.* U.S. Decennial Census.

into the reweighting procedure. By contrast, Richmond (Staten Island) exhibits minimal deviation throughout the entire period, consistent with relatively stable tract definitions. Kings (Brooklyn), New York (Manhattan), and the Bronx show moderate discrepancies, which decline over time as boundary stability improves. Notably, the deviation does not systematically bias population upward or downward but reflects localized distortions resulting from uniform distribution assumptions. Using EDs which are smaller than census tracts, increases precesion and leads to only minor deviations for the five boroughs. These deviations are particularly relevant when interpreting long-run neighborhood trends and estimating historical baseline conditions.

As discussed in Logan, Zhang, et al. (2021), spatial harmonization accuracy depends critically on the nature of boundary changes, with larger errors expected for non-nested, irregular transitions. While the area-weighting method provides a feasible approach to creating a tract-level panel, it may understate heterogeneity in neighborhoods with high internal variation or complex administrative histories.

C4 Newspaper Data

The rental price data used in this paper were collected from the real estate sections of the New York Times (NYT) for each decennial year between 1930 and 2010. The goal was to identify approximately 1,500 valid rental listings per sampled year. Sampling focused on the last Sunday of each month—when real estate listings were most abundant—across all 12 months. I sampled across all columns in the real estate section so that all areas covered by the newspaper would be included in the sample. Then, I randomly selected listings that satisfied three criteria: (1) the listing included a rental price, (2) it specified size (in rooms or bedrooms), and (3) it provided either the exact street address (with housenumber) or an intersection. If a sufficient number of qualifying listings could not be found for a given month, assistants consulted Sunday editions from adjacent months. An example of a typical NYT rental page is shown in Figure C4.1.

In the next step, addresses from NYT rental listings were geocoded using the Google Geocoding API. Listings were retained only if the geocoder successfully matched either an exact street address or a recognizable intersection (main and cross streets). Listings with incomplete or ambiguous location information were excluded. This procedure yielded a final sample of 19,446 rental listings across the eight decennial years. Table C4.1 summarizes rent levels from both U.S. Census tract data and NYT rental listings. Average rents increase steadily over time, with NYT-listed rents consistently exceeding census-based averages. This likely reflects the fact that NYT listings capture asking rents for advertised units — often skewed toward higher-end properties — whereas the census reports contract rents for existing tenants (i.e., sitting rents). Across all years, tracts with public housing

Figure C4.1: Example of a Rental Listings

8 W APARTMENTS Sunday advertisements must be ordered before 2 P. M. Saturday. THE

Apartments Furnished—Manhattan

Apartments of Three, Four, Five Rooms

Continued From Preceding Page

70'S (Park-Madison)—Sublet most unusual studio apartment, 4 rooms, 2 baths, beautifully furnished in the modern manner; dining gallery overlooks spacious duplex living room; cool, quiet. RHinelander 4-6437.

70'S, EAST—Three rooms, southern exposure, cross-ventilation; to October; reasonable. Sholl. Albert B. Ashforth, Inc., 12 E.44 MU.2-1100

70'S, Lower (off Central Park West)—4½ rooms; quiet; cool; beautifully furnished; 3 exposures; piano; May-October, longer; reasonable. ENdicott 2-2045.

70'S (lower, off Central Park West)—Sublet 4½ rooms till October, 1941; 3 exposures, spacious; $90. ENdicott 2-2045.

70'S, WEST—Unusual, beautiful, large 3 rooms; reasonable right party. TRafalgar 7-6848.

70'S, WEST—4 (2 Bedrooms), Elegantly Furnished! Outstanding Bargain; $80. TRafalgar 4-3156.

71ST (240 West End) (14B)—Three large, cool rooms, kitchen, exquisite, completely furnished; accommodating 2 or 4 persons.

71ST, 325 WEST—Living room, bedroom, complete kitchen, bath; elevator; telephone; reasonable.

72D, EAST—Sublet charmingly furnished 3½ (16th floor); outside dinette-kitchen; unobstructed; cross-ventilation; October-longer. RHinelander 4-8785.

72D, 405 EAST—3 large rooms; May 15 occupancy. RE. 7-4772, afternoons or evenings; also unfurnished.

72D, EAST—4½ rooms, 2 baths, June-Oct., $100. BUtterfield 8-6882, Monday.

72D, 124 WEST—Luxurious 3 rooms, 3 exposures, surrounding terrace, $125. Supt.

73D, 28 EAST—Summer bargain. Attractively furnished, 3 large room, bath, kitchen, dinette, extra toilet; cross draft; new elevator building. Waterman, ELdorado 5-6900.

73D, 210 EAST (9C)—Furnished 4 rooms, 2 baths; 9th floor; East Village; sublet Summer. Kahn. BUtterfield 8-3886.

73D, 105 WEST—Modern elevator, 3 rooms, large kitchen, unobstructed view; lease, $65. SU. 7-5554, SA.

73D, 40 WEST—Airy 3-room apartment, attractively furnished, bath, kitchenette, frigidaire.

75TH ST., 157 EAST—Cooperative building; living room, bedroom, dressing room and bath; fireplace; well furnished; $100; excellent meals served if desired; selective tenantry. Apply Mrs. Brown, premises. RHinelander 4-0550.

Apartments Furnished—Manhattan

Apartments of Three, Four, Five Rooms

187TH, 825 WEST (overlooking Hudson River—Sublet 4 rooms, furnished, with antiques, unusual view, reasonable rental. See Supt.

189TH, 701 WEST (Broadway) (5E)—Sublet, 3½ modern rooms, cozy, elevator; $55.

BEEKMAN PLACE

Attractive Summer Rental

4½ rooms, master bedroom, paneled nursery or library, connecting baths, large living room, dining foyer, kitchen, cross-ventilation; East River view; immediate occupancy until October, by appointment. Telephone PLaza 8-0349. REASONABLE RENTAL.

BEEKMAN PLACE, 32—Sublease now until October, attractive 3 rooms, cool, smart, river view. Sacrifice. Call PLaza 5-3155.

CABRINI BLVD., 160 (Castle Village) (Apt. 36)—Artist's beautifully decorated 4½-room apartment facing Hudson; sublet furnished-unfurnished; immediate occupancy to October. WAdsworth 7-2816.

CENTRAL PARK WEST (1 West 85th)—Five rooms, 7 windows overlooking park; sacrifice; piano, radio; complete. TRafalgar 7-5321.

CENTRAL PARK WEST—Attractive 4 light rooms, 2 baths, cross ventilation; June-October; references. TRafalgar 7-4378.

CENTRAL PARK WEST, 25—5 rooms to sublet till October; attractive furnishings. Apply renting office, Century Apartments.

CLAREMONT AVE., 191 (Columbia University)—Delightful 4 rooms, elevator, piano; $58. Hankins.

EAST END AVE.—Sublet to October; 3 rooms, bath, kitchen, cross-ventilation; top floor, walk-up; worth it for gorgeous view across Carl Schurz Park and miles of East River; also roof privileges; $50. REgent 7-4152.

GRAMERCY SECTION, modern building, 3½ rooms, beautifully furnished, roof garden, maid service; 15th floor; sacrifice. GRamercy 3-4667, or office, STuyvesant 9-3884.

GRAMERCY PARK—Unusually attractive apartment; beautifully furnished 5 rooms, magnificent view; 3 exposures; 12th floor; $150. GRamercy 5-4029.

GRAMERCY PARK, 8—3½ rooms, new; 3 exposures; all housekeeping appliances, radio; 4 months lease. Supt.

GRAMERCY PARK NORTH, 60—3-4, modern, twin beds, dinette. GRamercy 3-7399.

GREENWICH VILLAGE—Comfortable, attractive sublet, elevator building, 3 rooms, 2 exposures; $65.

Apartments Furnished—Manhattan

Apartments of Six Rooms and Over

RIVERSIDE DRIVE in 80's: 7 outside rooms, 3 baths, 10th floor, marvelous views, grand piano, 4 radios, television; rent for Summer or longer; $185 per month. TRafalgar 4-0208.

RIVERSIDE DRIVE—Sacrifice 6 rooms, new Steinway grand; river view. Algra Realty Corp., 70 Pine St.

FURNISHED beautiful Park Ave. 6-room apartment, real home, 3 baths, 35-ft. living room, den, dining, kitchen, 2 master bedrooms, maid's room; best transportation everywhere; for 3 to 5 months to responsible party; $300. W 486 Times.

EXCEPTIONALLY well-furnished apartment, 6 rooms, all 4 exposures; $150 monthly until October or longer. SUsquehanna 7-1745.

MUSICIAN sublet furnished 6 rooms, grand piano; beautiful river view. MOnument 2-3726.

Penthouse—Terrace Apartments

Furnished

1ST TO 5TH AVES.—EAST-WEST.
Mrs. Ida Catlin, 1049 Lexington Ave.
PENTHOUSE SPECIALIST.
Unfurnished-furnished and apartments; many attractive offerings; excellent values; mail orders; state requirements; immediate or October. REgent 4-1400.

5TH AVE.—24TH FLOOR—DUPLEX
40' Salon, 4 Masters, 5 Baths. Landscaped Terraces. Also Unfurnished. BUtterfield 8-9430.

38TH, 137 EAST—Sublet approximately May 25, Sept. 25; unusually distinctive 3½ rooms, modern penthouse; large terrace; reasonable to suitable tenant. CAledonia 5-5360.

40TH (between Madison and 5th)—Deluxe duplex penthouse apartment; living room, dining room, kitchen and pantry; maid's room and bath; 2 master bedrooms, 3 baths, terrace; handsome antique furnishings; $600 month; occupancy April 30. Call for appointment. MUrray Hill 3-5245.

50'S, EAST—Unusual penthouse; studio living room, 2 master bedrooms, 3 baths, dining room, bar, kitchen, pantry, maid's room, fireplaces; terrace surrounds apartment; furnished by decorator; also unfurnished. PLaza 3-5058.

50'S, EAST—River view; 5-6 room duplex, with roof terrace, large living room, separated by plate glass partition from dining room; grand piano; furnished-unfurnished. Weekdays, Chace, PLaza 3-1700.

50'S (Sutton Place Vicinity)—River View, Garden Living Room; three rooms, bath; exquisite furnishings. $135. Immediate. PLaza 3-5572.

Note. A sample real estate listing from the NY Times in 1940.
Source. NYT 28.04.1940.

exhibit lower average rents, confirming the dampening effect of public housing on local price levels. While total rents rise sharply across decades, rent per room increases more gradually, suggesting improvements in unit size or quality. Among all measures, the strongest cross-source similarity is found between census average rents and NYT median rent per room. This suggests that per-room measures from NYT listings provide the most comparable counterpart to tract-level census rents. Finally, the number of listings is largest in 1930, reflecting intentional oversampling to ensure strong baseline coverage across tracts.

Table C4.1: Rent Price Statistics

	US census		NYT rental listings					
Year	Avg. rent	Avg. rent no p.h.	Med. Rent	Med rent p.r.	Avg. rent	Avg. rent p.r.	Avg rooms	Obs
1930	808.22	836.31	1358.23	507.88	1877.25	712.47	3.63	8847
	(2505.53)	(2637.84)			(1867.09)	(771.37)	(2.51)	
1940	605.45	618.01	1168.39	389.46	1840.45	615.18	3.64	1561
	(185.39)	(185.94)			(2386)	(1121.86)	(2.28)	
1950	342.83	346.55	1076.93	434.39	1332.28	511.38	2.99	1572
	(174.18)	(179.16)			(1408.2)	(357.42)	(1.51)	
1960	431.71	441.39	1053.66	324.20	1289.63	390.17	3.46	1632
	(164.9)	(167.46)			(756.11)	(187.73)	(1.34)	
1970	574.78	587.08	1826.87	615.51	2058.89	774.6	3.37	1432
	(226.33)	(233.7)			(1153.69)	(557.55)	(1.72)	
1980	576.6	593.11	1574.87	741.12	1886.81	1000.31	2.81	1509
	(174)	(173)			(1437.99)	(844.8)	(1.98)	
1990	806.78	839.85	1919.01	1251.53	2212.58	1302.05	2.48	1585
	(237.28)	(223.46)			(1218.39)	(796.87)	(2.12)	
2000	864.51	901.53	3293.03	1618.02	3819.1	1900.45	2.89	1197
	(304.1)	(293.09)			(2586.1)	(1224.96)	(2.07)	
2010	1132.89	1184.11	2700.00	1297.50	3483.18	1612.85	3.24	111
	(339.46)	(308.38)			(2584.45)	(1161.03)	(2.41)	

Note. Table C4.1 Figure C4.1 presents descriptive statistics for rent prices using data from the U.S. Census and New York Times (NYT) rental listings, with all prices deflated to 2010 dollars. Columns (2) and (3) report the average of census tract-level median contract rents; column (3) excludes tracts with public housing. Columns (4) and (5) show the median rent and median rent per room from NYT listings. Columns (6) and (7) report the average rent and average rent per room. Column (8) gives the average number of rooms per listing, and column (9) shows the number of listings per year. Standard deviations are reported in parentheses.

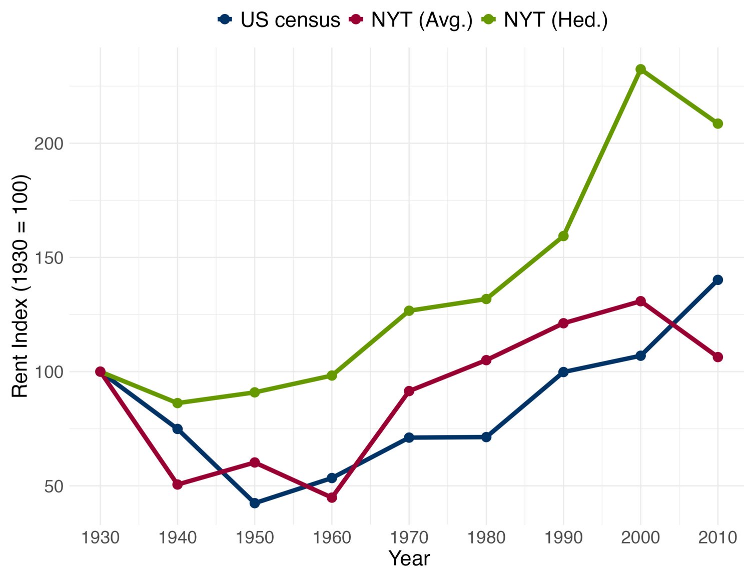
While Table C4.1 compares raw levels, I further assess the consistency of rental price trends across data sources by constructing rent indices from NYT listings and comparing them to an index based on U.S. Census tract-level median contract rents. To ensure comparability, all series are normalized to 1930 (=100). Because the NYT listings reflect advertised asking rents, which may vary by unit quality, location, and timing, I adjust for compositional changes over time including unit characteristics such as included utilities, furnishings, and parking availability, along with census tract and seasonal fixed effects. I estimate the following hedonic pricing model that adjusts for observable unit characteristics and location-specific factors:

$$\log(rent_{ikmt}) = \alpha + \sum_{t=1940}^{2010} \beta_t \mathbb{1}(t = year) + \mathbf{X}_i + \gamma_r + \gamma_k + \gamma_m + \epsilon_{ikmt} \tag{C.1}$$

The coefficients β_t capture changes in average advertised rents over time, relative to 1930. The vector of controls $\mathbf{X}_i$ includes indicators for unit characteristics that were consistently reported in the listings: whether gas, water, electricity, or air conditioning were included in rent; whether parking was available; and whether the unit was furnished. The specification also includes room fixed effects γ_r, census tract fixed effects γ_k, and calendar month fixed effects γ_m to account for differences in unit size, location, and seasonality.

Figure C4.2 displays the resulting rent indices. The census-based rent index shows a 45 percentage point decline until 1950, followed by a steady increase through 2000 and a slight uptick in 2010. The average rent per room index from NYT listings closely tracks this trend—declining in the 1940s and 1950s, then rising in subsequent decades. The NYT-based hedonic index exhibits a similar trajectory: a modest dip in the early period, strong growth through 2000, and a modest decline in 2010. Most importantly, the trends across both data sources run nearly parallel from 1950 to 2000. This suggests that, even after accounting for sample differences, both NYT listings and census rents provide a reliable signal of long-run rent dynamics.

Figure C4.2: Rent Indeces



Note. Figure C4.2 plots the rent indices from two sources. The blue line shows the average of census tract-level median contract rents. The rent line shows the average of rent per room from the New York Times listings. The green line shows the β_t coefficients from the hedonic regression in Equation C.1, estimated using NYT rental listings and controlling for property characteristics, census tract fixed effects, and month-of-year dummies. Both series are indexed to 1930 (=100) for comparability. All rent prices were deflated to 2010 dollars using the CPI.

D Extended Results

D1 Hedonic rent measure

To capture changes in neighborhood rent levels net of compositional changes in unit quality, I construct quality-adjusted (hedonic) rent estimates at the census tract-by-year level using historical rental listings from the New York Times. The analysis proceeds in two steps.

In the first step, I estimate a hedonic regression of log monthly rent on observable unit characteristics reported in the listings. These include indicators for whether utilities (electricity, water, gas) are included, whether the unit is furnished, whether parking is available, and the number of rooms. Given New York City's long-standing rent regulation regime—initiated under federal rent ceilings in 1943 and continued through local rent control and stabilization laws — I also include indicators for whether a listing was rent - controlled or stabilized, when such information was provided by the advertiser. All monthly rent values are deflated to 2010 dollars using the Consumer Price Index (CPI), ensuring comparability across years. The estimating equation is:

$$\log(rent_{ikrmt}) = \alpha + \mathbf{X}_i + \iota_{kt} + \gamma_r + \gamma_m + \epsilon_{ikrmt} \tag{D.1}$$

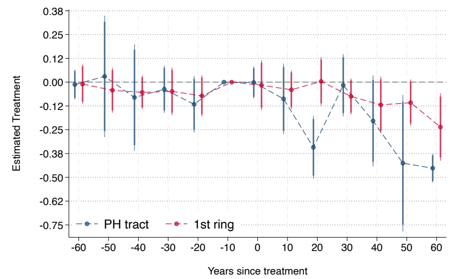
Here, $\mathbf{X}i$ includes the listing-specific characteristics described above; γ_r and γ_m are fixed effects for the number of rooms and calendar month, respectively; and ι_{kt} is a full set of census tract-by-year fixed effects. These fixed effects capture residual variation in rent levels across space and time, controlling for unit quality and seasonality.

In the second step, I use the estimated ι_{kt} coefficients as the dependent variable in the event study framework described in Equation 2. This allows me to estimate how public housing construction influenced neighborhood rent levels after adjusting for observable differences in rental unit quality and macroeconomic inflation. Unlike the outcomes obtained from the census — namely, the White, Black, and Hispanic population and the median contract rent — I do not control for the baseline level of the hedonic rent outcome. This is because the rental listing data do not consistently cover all census tracts in all years, and I do not observe a full panel of ι_{kt} estimates across time. Results are presented in Figure D1.1.

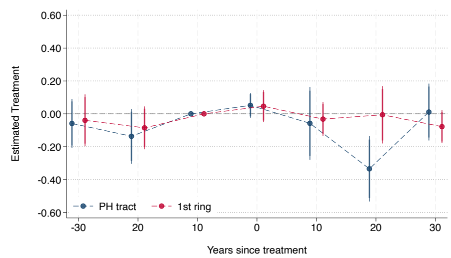
Panel a shows the average treatment effect across all projects. By 40 to 60 years post-construction, rents in treated areas are 18.37% and 36.36% below baseline levels respectively. Importantly, the estimated effect at the 30-year mark appears smaller in magnitude (-1.59%) before declining again thereafter. This local deviation from the long-run trend is not indicative of treatment reversal, but more likely reflects selective reporting in the newspaper rental listings. In particular, if higher-quality units are more likely to be advertised, the rent distribution captured in the data may temporarily shift upward, mechanically attenuating the treatment estimate in that period.

Figure D1.1: Effect of Public Housing on Hedonic Rent

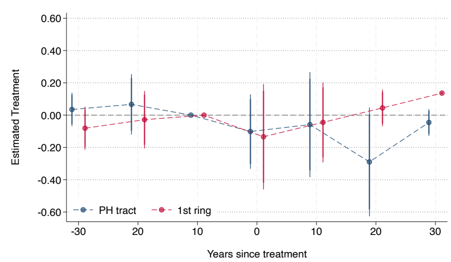
(a) log(Hedonic Rent)



(b) log(Hedonic Rent) - pre 1970



(c) log(Hedonic Rent) - pos 1970



Note: Figure D1.1 plots coefficients $\hat{\beta}_\tau$ from Equation 2, where the outcome variable is the log of quality-adjusted rent obtained from a hedonic regression. Panel a uses the full sample of public housing projects. Panel b restricts the sample to projects constructed before 1970, while Panel c includes only those built after 1970. Standard errors are clustered at the census tract level. Vertical lines indicate 90% and 95% confidence intervals. The omitted category corresponds to census tracts located in the second ring around each project.

Panel b restricts the sample to projects constructed before 1970 and reveals a pattern consistent with the full sample: quality-adjusted rents in treated tracts decline steadily, reaching a trough of approximately –28.4% twenty years after construction. However, by year 30, there is a notable rebound to a near-zero effect (1.1%). This mid-horizon bounce-back likely reflects selection in the newspaper listings rather than a true recovery, as shifts in the composition of advertised units may bias average rent estimates upward over time.

Importantly, there is no corresponding movement in adjacent tracts, indicating that the reversal is not driven by neighborhood-wide shifts.

Panel c focuses on projects constructed after 1970. While coeffecicents are insiganicant they reflect a pattern more constente with a moderate effect. Rent fall by around 7.5% and -25.17% 30 years and bounce back to -4%. While there is clearly no positive effect of rent prices, the effect on market rents of later constructed projects seems week.

D2 Heterogeneity Analysis

This section investigates whether the effects of public housing construction vary systematically across space and socioeconomic status. I examine heterogeneity along two key dimensions: (i) borough of New York City and (ii) neighborhood income level. To do so, I extend the baseline stacked event study specification by interacting treatment indicators with dummies for each borough and for income tertiles, defined at the neighborhood (NTA) level.

- 1st tertile (lowest): \$15,559–\$29,668
- 2nd tertile (middle): \$29,673–\$41,227
- 3rd tertile (highest): \$41,228–\$78,009

Income values are based on median household income in the pre-treatment year and have been deflated using the 2010 CPI to ensure comparability across time. The estimation follows a triple-indexed difference-in-differences framework:

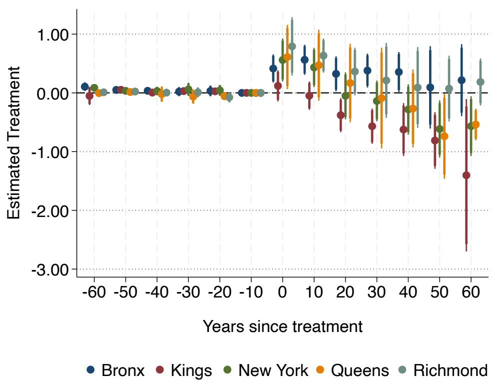
$$y_{k,m,p,t} = \sum_{h \in H} \sum_{R} \sum_{\tau=-60}^{60} \beta^{(h)}_{\tau,r} \cdot \mathbb{1}\left(t - Y_p = \tau, R = R(k,p), h = h(m)\right) + \mathbf{X}_{p,k} + \rho_{p,t} + \rho_{p,m,t} + u_{k,m,p,t} \tag{D.2}$$

where $h \in H$ indexes either the borough or income group, and r indicates proximity ring (e.g. public housing tract or first ring). This setup allows for estimating differential dynamic treatment effects across subgroups. To construct income groupings, I use tract-level real median household income from the decade before project completion. I average tract incomes to the NTA level and assign each NTA to an income tertile based on its position in the citywide distribution. These income groups are used to stratify treatment effects.

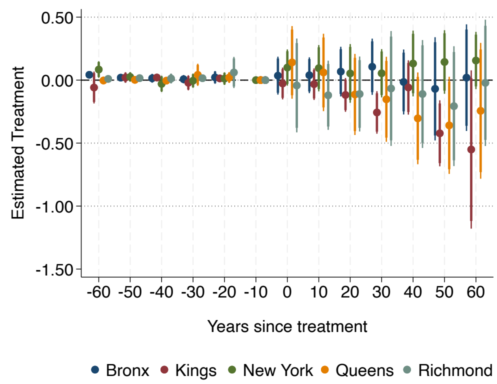
Borough-Level Heterogeneity. The borough-specific results (Figure D2.1) show consistent qualitative patterns across the five boroughs. In each case, public housing led to long-run declines in White population and sustained increases in Black population within treated tracts. These demographic changes begin shortly after project completion and persist for several decades. Effects are most pronounced in the Bronx and Manhattan but are directionally similar in Brooklyn, Queens, and Staten Island. Rent responses are more variable: some boroughs show short-run increases followed by attenuation, others display flat or slightly negative trends. Importantly, no borough exhibits a reversal or absence of effect, suggesting robust and spatially consistent treatment effects.

Figure D2.1: Effect by Borough

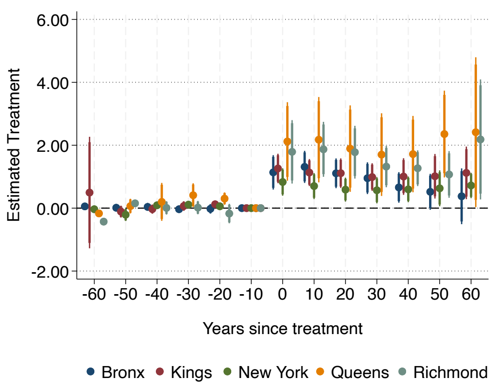
(a) log(White) - PH tract



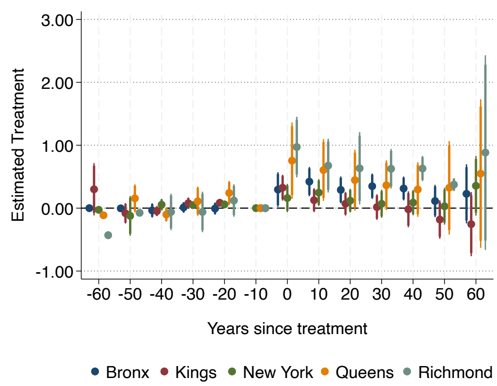
(b) log(White) - 1st ring



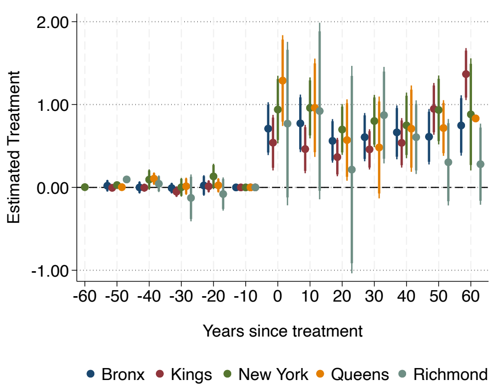
(c) log(Black) - PH tract



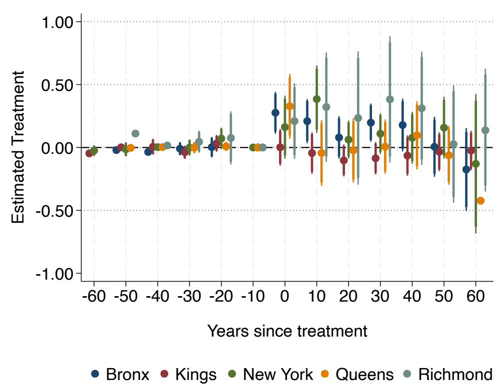
(d) log(Black) - 1st ring

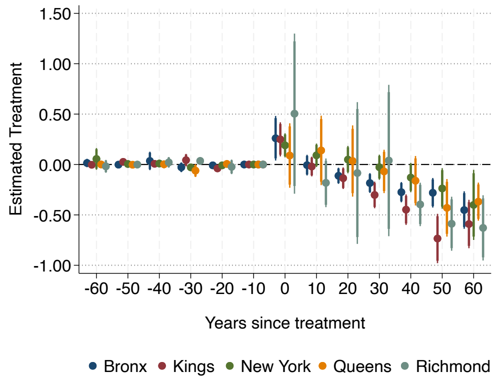


(e) log(Hispanic) - PH tract

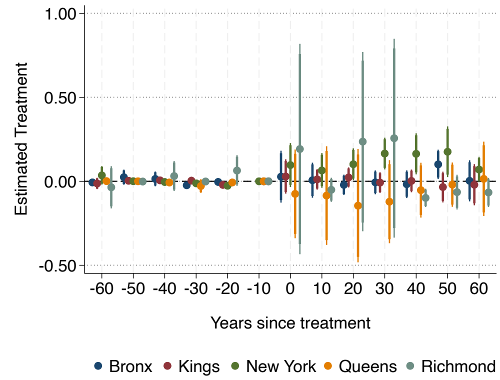


(f) log(Hispanic) - 1st ring





(g) log(Rent) – PH tract



(h) log(Rent) – 1st ring

Note: Figure D2.1 plots report coefficients $\hat{\beta}_{\tau,r}$ in Equation D.2 for each treated tract and rings around a project; Panels a, c, e and g show effects for census tracts containing a public housing project (PH tract). Panels b, d, f and h show effects for adjacent tracts in the first ring (1st ring).coefficients have been interacted with an indicator variable for each of the five boroughs in New York: Bronx (blue), Kings (red), New York (green), Queens (orange), and Richmond (teal); standard errors are clustered at the project level; the vertical lines show the estimated 90% and 95% confidence intervals; the omitted category consists of tracts within a second ring in each borough. Panels a through f use weighted unit counts from the US Census; estimates have been weighted by frequency by ring.

Income-Based Heterogeneity. Figure D2.2 displays results by income tertile. Across all income groups, public housing leads to a sharp and persistent decline in the White population and a corresponding increase in the Black population in treated tracts, with smaller but similar patterns in adjacent areas. While low-income areas (1st tertile) exhibit somewhat larger demographic responses, the direction and timing of effects are consistent across tertiles. Rent effects are also broadly similar: in all income groups, rents initially increase following construction, but either plateau or decline modestly in the long run. The only exception is a slight upward pressure observed in adjacent tracts (first ring) of high-income areas (3rd tertile), where long-run rent effects remain mildly positive. There is no income group in which the treatment effect reverses or disappears, suggesting a high degree of homogeneity in response to public housing across the income distribution.

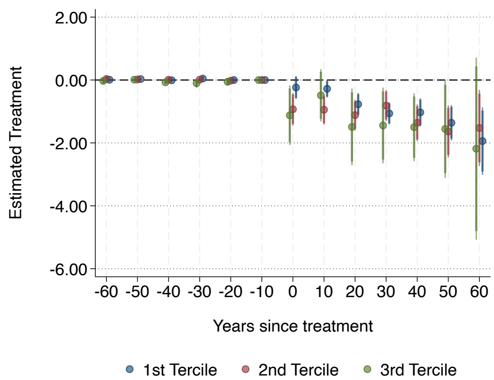
D3 Excluding Overlapping Rings

In this section, I restrict the analysis to a subset of public housing projects with non-overlapping treatment and control rings. Specifically, I drop all sub-experiments where the first or second ring of a given project overlaps with the treatment or spillover area of another project. This ensures that no tract is simultaneously assigned to multiple roles (e.g., treated and control) across different sub-experiments. The resulting sample consists of 18 projects constructed between 1950 and 1973 that satisfy this non-overlap condition.

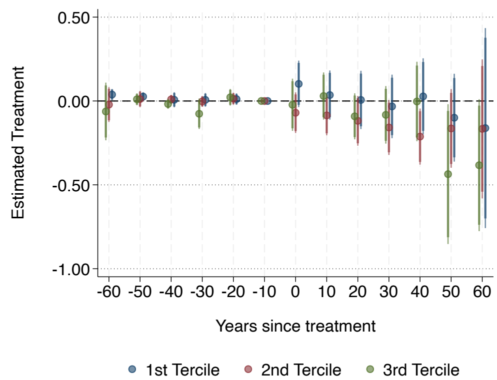
Restricting to non-overlapping control rings improves internal validity by eliminating

Figure D2.2: Effect by Income Tertile

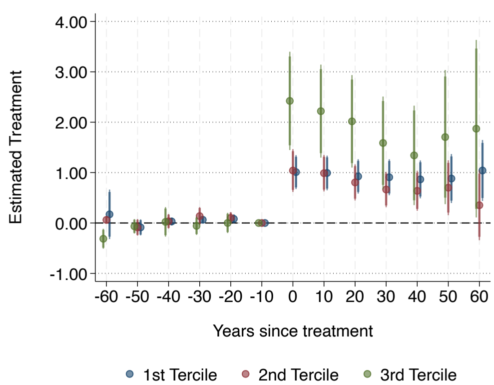
(a) log(White) - PH tract



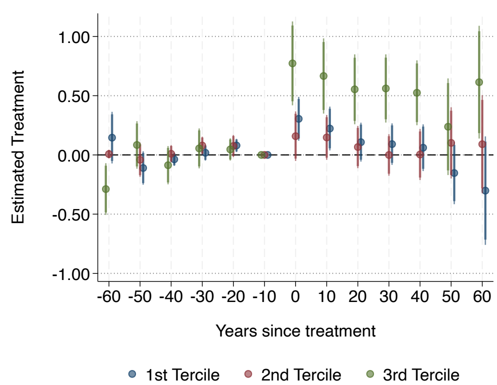
(b) log(White) - 1st ring



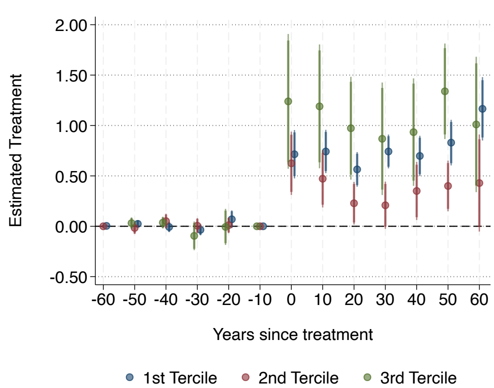
(c) log(Black) - PH tract



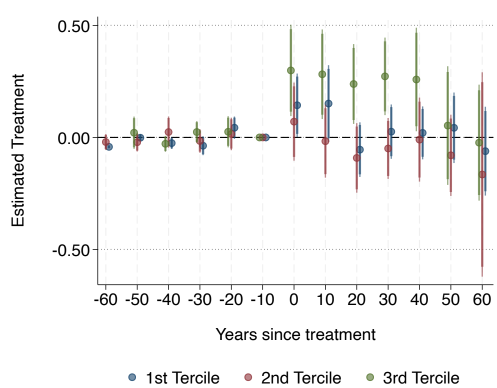
(d) log(Black) - 1st ring

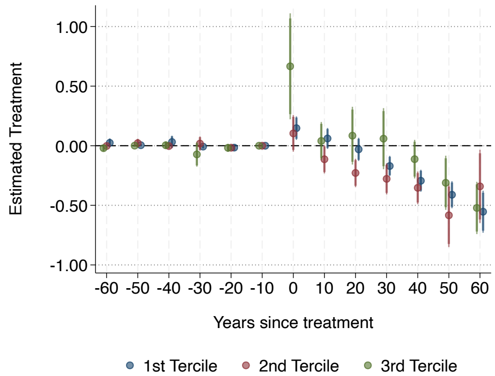


(e) log(Hispanics) - PH tract

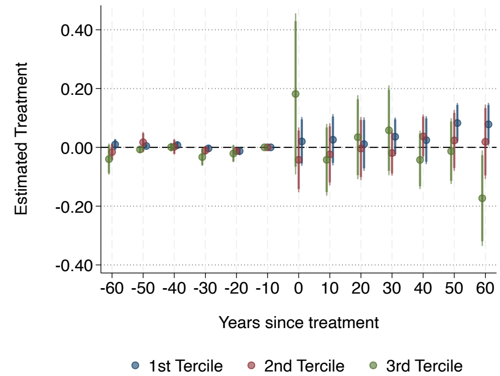


(f) log(Hispanics) - 1st ring





(g) log(Rent) - PH tract



(h) log(Rent) - 1st ring

Note: Figure D2.2 plots report coefficients $\hat{\beta}_{\tau,r}$ from Equation D.2 for each event time τ by treatment proximity and income tertile. Panels a, c, e and g show effects for census tracts containing a public housing project (PH tract). Panels b, d, f and h show effects for adjacent tracts in the first ring (1st ring). Each coefficient is interacted with an indicator for income tertile based on NTA-level pre-treatment income. Standard errors are clustered at the project level; vertical lines show 90% and 95% confidence intervals. All estimates are weighted by tract frequency within ring and income group.

potential spillover from contamination of control areas. Overlap could bias estimates by introducing indirect treatment exposure in the comparison group, violating the identifying assumption that control tracts remain unaffected by nearby treatments. I re-estimate Equation (D.2) using only this restricted set of projects. This restriction helps isolate the dynamic effects of early public housing without interference from later projects. Doing so gives 13 subexpeimnets with non overlapping rings, constrcuted between 1940 and 1984.[20]

Figure D3.1 presents the long-run effects of public housing construction using the restricted sample of 18 projects with non-overlapping treatment and control rings. Panel a shows a persistent and substantial decline in the net White population in treated tracts—falling by 90.34% thirty years after construction and by 97.61% after sixty years. This implies an almost complete exit of White households living in private housing within these areas. In contrast, the total White population, which includes public housing residents, declines less sharply and is not consistently statistically significant, suggesting that some White residents continued to enter or remain in public housing developments. There is no evidence of similar declines in adjacent areas, indicating that these changes were highly localized and not the result of broader displacement.

Panel b reports the trajectory of the Black population. Immediately following construction, the total Black population in treated tracts rises sharply by 286.90%, reflecting an influx into public housing. However, over time, the net Black population—excluding public

[20]The project names are: Bland, Carey Gardens, Clason Point Gardens, Fenimore-Lefferts, Fulton, Glenwood, Hammel , Ingersoll, La Guardia, Mitchel, Morrisania Air Rights, New Lane Area, Nostrand, Ocean Bay Apartments (Oceanside), Riis II, Robbins Plaza, South Jamaica I, Surfside Gardens.

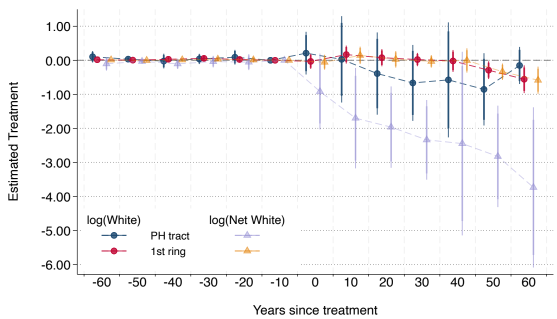
housing residents—declines significantly. Thirty years post-construction, the private Black population falls by 62.28%, reaching a 91.03% decline by year sixty. As with Whites, these effects are confined to treated tracts, with no comparable changes in adjacent neighborhoods.

Panel c shows more muted patterns for the Hispanic population. The net Hispanic population remains relatively flat and statistically insignificant throughout the post-treatment period, suggesting little change among privately housed Hispanic residents. In contrast, the total Hispanic population increases by 142.37% on average, indicating that the observed growth was driven almost entirely by increases in the public housing population.

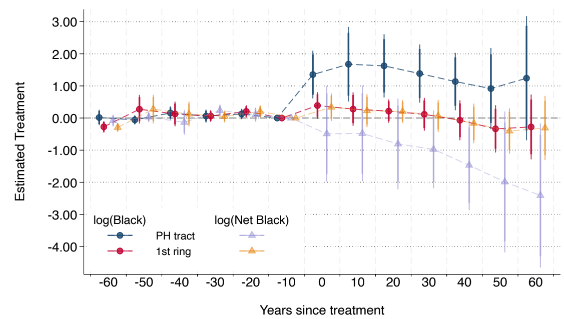
Panel d displays effects on median contract rents using Census data. Rents in treated tracts decline steadily over time, falling by 24.95% after forty years, while rents in adjacent tracts remain stable and statistically indistinguishable from zero.

Figure D3.1: Long-run effects public housing (non overlapping)

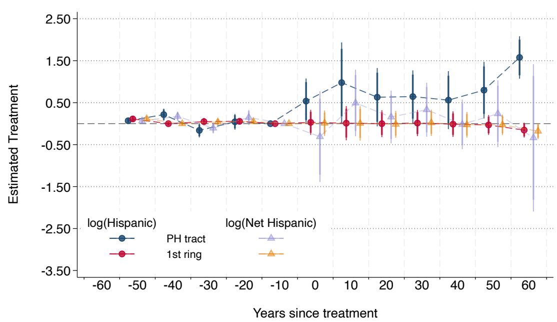
(a) White



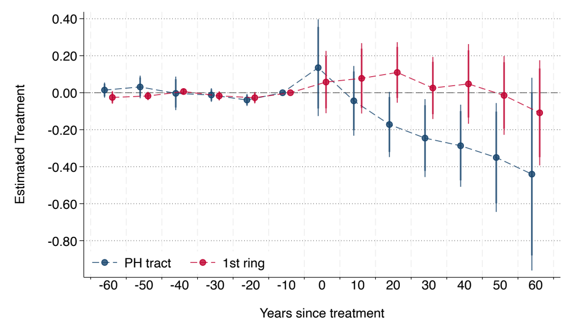
(b) Black



(c) Hispanics



(d) log(Rent)



Note: Figure D3.1 displays coefficients $\hat{\beta}_{\tau,r}$ from Equation 2, estimated on a restricted sample of 18 public housing projects with non-overlapping treatment and control rings. Panels a to c use as outcomes the log of net White, Black, and Hispanic population — that is, total population of each group excluding public housing residents. Panel d reports results for the log of median contract rent.
Colors indicate the ring of exposure: blue and purple for treated tracts (within-project); red and orange for adjacent tracts (first ring). The omitted category consists of tracts in the second ring. All estimates are weighted by tract frequency within each ring, and standard errors are clustered at the project level. Vertical lines show 90% and 95% confidence intervals.

D4 Event Study Results - Panel Setup

In this Section, I report event study results in this section using alternative estimators that correct for the shortcomings of standard two-way fixed-effects (TWFE) models. Specifically, the literature focused on the "forbidden" comparison between later-treated and earlier-treated units, which the TWFE estimator might not handle correctly. As shown in Goodman-Bacon (2021), the TWFE estimator might choose weights that lead to the estimator having the wrong sign. The estimators proposed in the literature differ in terms of who they use as the comparison group (e.g., not-yet-treated versus never-treated) and the pre-treatment periods used in the comparisons (e.g., the entire pre-treatment period versus the final untreated period).[21]

To test the coherence of the approach using a stacked design, as proposed in Section 4, I use the panel setup. In this setup, a tract is treated when it has had a public housing project within its boundaries at any point in time. To serve as the appropriate control group, I compare treated tracts to tracts in the second ring, surrounding the inner ring. This is motivated by two reasons. First, the second ring serves as a coherent control group from the stacked to the panel setup. Second, since it is reasonable to assume public housing generates spillovers, dropping the first tract ring around public housing will suffice not to violate SUTVA. Figure B.1 shows the spatial layout of treatment and control. I estimate the following dynamic specification:

$$y_{k,m,t} = \sum_{\tau=-60}^{70} \beta_\tau \left(t - Y_p\right) + \mathbf{X}_i + \rho_t + \zeta_k + \Xi_{m,t} + u_{k,m,t} \tag{D.3}$$

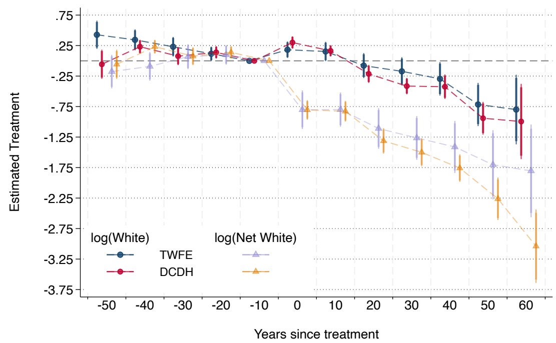
The parameter of interest, denoted as β_τ, captures the effect of the arrival of public housing in census year t relative to the year of construction Y_p compared to the outermost rings. I control for census year ρ_t and tract ζ_i fixed effects. Finally, I allow tracts within a neighborhood to trend differently each year by including non-parametric neighborhood trends $\Xi_{m,t}$. Finally I account for local infrastructure developments that may independently affect neighborhood outcomes, ($\mathbf{X}_i$). Specifically, I control for whether a tract was exposed to (i) an urban renewal project, (ii) an arterial road expansion, or (iii) a new subway connection, using time-varying treatment indicators. Results from estimating Section D.3 are shown in Figure D4.1.

Figure D4.1 displays dynamic treatment effect estimates from two different estimators: a standard Two-Way Fixed Effects (TWFE) model and the estimator proposed by De Chaisemartin and D'Haultfœuille (2020) (DCDH). Panel a shows the estimated effect on the White population. Under the TWFE specification, the long-run effect appears attenuated

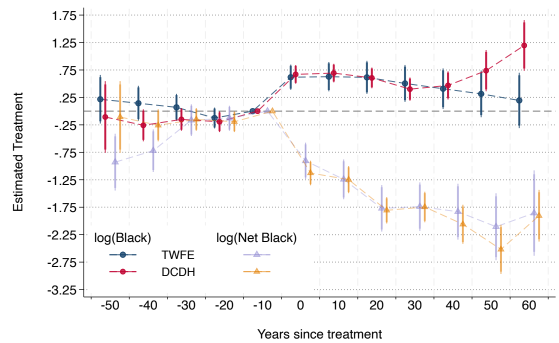
[21] I refer to Roth et al. (2023) for an excellent overview of recent advancements in the DiD literature and practical guidance on how these estimators differ.

Figure D4.1: Effect of Public Housing

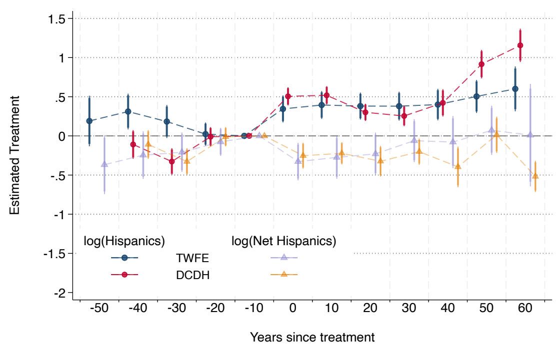
(a) White



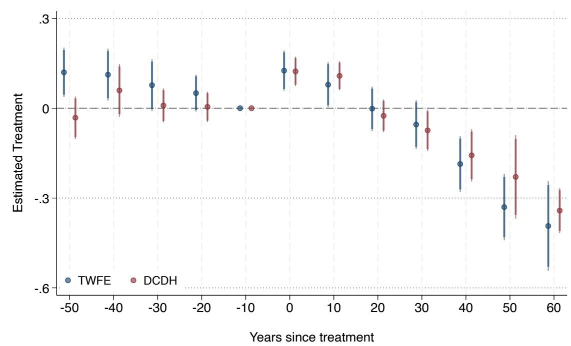
(b) Black



(c) Hispanics



(d) log(Rent)



Note. Figure D4.1 displays event study coefficients $\hat{\beta}_\tau$ estimated from Equation D.3. Panels a, b, and c report results for White, Black, and Hispanic populations, respectively, while Panel d shows effects on median contract rent (in 2010 dollars). For details on outcome variables, see Section 3. In Panels a–c, blue and red lines correspond to TWFE and DCDH estimators on total population, while purple and orange lines represent TWFE and DCDH estimates net of public housing residents (i.e., private population). Estimators are abbreviated as follows: TWFE = Two-Way Fixed Effects; DCDH = de Chaisemartin and D'Haultfoeuille estimator (De Chaisemartin and D'Haultfœuille, 2020). Vertical lines indicate 90% and 95% confidence intervals. Standard errors are clustered at the census tract level.

and less precisely estimated. In contrast, the DCDH estimator shows a more pronounced and persistent decline, suggesting stronger long-run effects. Panels b and c show corresponding effects on Black and Hispanic population. Here, both estimators detect sustained post-treatment increases, but again, the DCDH estimates are larger and more stable. Panel d reports effects on log median contract rents. The TWFE estimator indicates modest increases in the short run followed by attenuation, while the DCDH estimator finds more persistent and slightly stronger effects.

The DCDH estimator improves upon TWFE by avoiding "forbidden comparisons" — specifically, it does not use already-treated units as controls for later-treated units. This correction addresses bias in TWFE when treatment timing is staggered and effects are heterogeneous over time. As a result, the DCDH estimates likely better reflect the true long-run impact of public housing construction. These treatment effects are highly consistent with the long-run estimates reported in Figure 2.

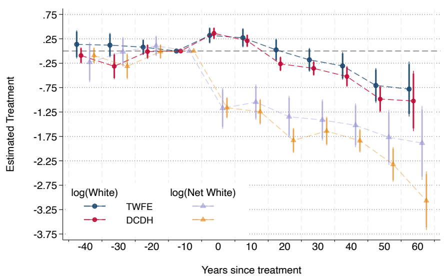
Figure D4.2 presents dynamic treatment effects separately for projects constructed before and after 1970. Panels a and b show effects on the White population. For pre-1970 projects, there is a sharp and persistent decline in White residents beginning shortly after project completion. In contrast, post-1970 projects have much smaller and statistically weaker effects. Panels c and d show similarly divergent patterns for Black population: strong and sustained increases following pre-1970 projects, but muted or temporary effects for those built later. Rent effects in Panels e and f follow the same structure—initial increases and long-run declines for older projects, but flat or null effects post-1970.

Notably, for projects constructed after 1970, the estimated effects on the Black population differ from those in Figure 4, where the control group is more finely defined using spatial rings. In the panel setup, the comparison group consists of all tracts in the second ring, which may introduce noise due to broader heterogeneity in untreated areas.

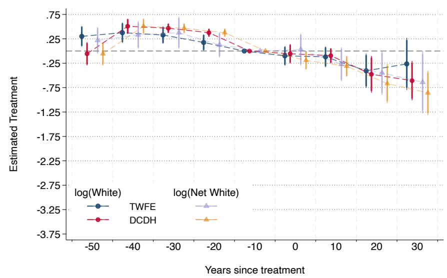
These results reinforce the conclusion that early public housing projects had large and persistent neighborhood effects, while later projects—likely due to changes in policy, design, or public sentiment—had a much weaker footprint.

Figure D4.2: Effect of Public Housing before and after 1970

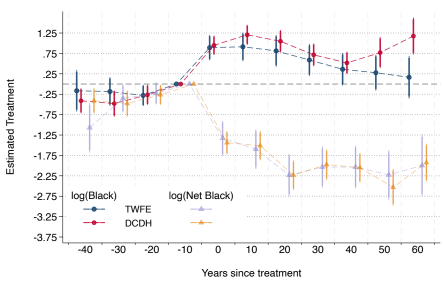
(a) White - pre 1970



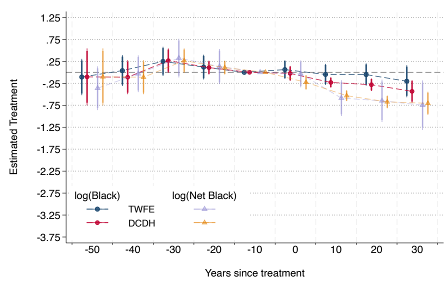
(b) White - post 1970



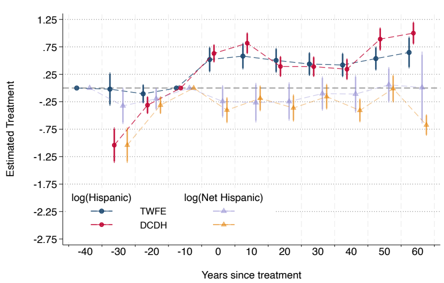
(c) Black - pre 1970



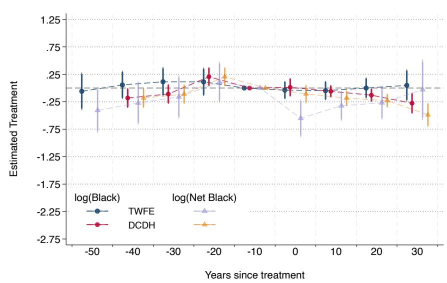
(d) Black - post 1970

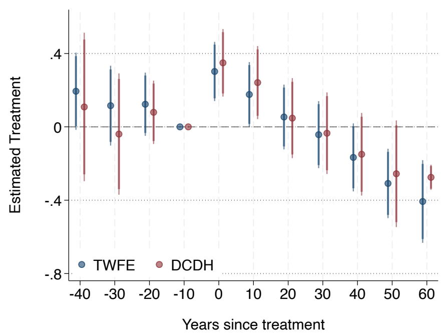


(e) Hispanic - pre 1970

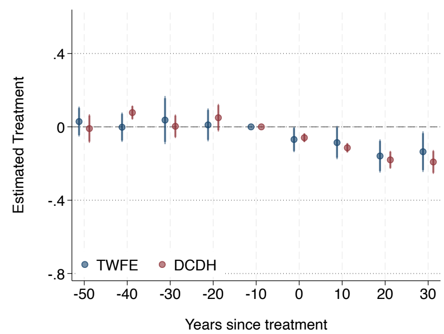


(f) Hispanic - post 1970





(g) log(Rent) - pre 1970



(h) log(Rent) - post 1970

Note: Figure D4.2 plots coefficients $\hat{\beta}_\tau$ from Equation (D.3), estimated using both TWFE and DCDH estimators. For the TWFE specification, treatment effects are estimated jointly by interacting event-time indicators (leads and lags) with a dummy for whether the project was constructed before or after 1970. For the DCDH estimator, effects are estimated separately by sample: pre-1970 effects use only early projects, and post-1970 effects use only later projects. Each panel reports estimates for a specific outcome: Panels a, c, and e show results for pre-1970 projects; Panels b, d, and f for post-1970 projects. In Panels a–f, blue and red lines correspond to TWFE and DCDH estimates on total population, while purple and orange lines show TWFE and DCDH estimates net of public housing residents (i.e., private population). Panels g and h show effects on median contract rent. Standard errors are clustered at the census tract level, and vertical lines indicate 90% and 95% confidence intervals. The omitted category consists of tracts in the second ring.

D5 MWTP for post-1970 Projects

In this section, I estimate marginal willingness to pay (MWTP) for public housing projects constructed after 1970, using the same empirical strategy described in Section 8. Specifically, I rely on reduced-form estimates from Equation (10), where treatment effects are interacted with time-invariant public housing attributes. The analysis is limited to a 0–30 year post-construction horizon, as the earliest post-1970 projects in the sample were completed around 1980, providing at most three decades of post-treatment data. The outcome variables used in this analysis are the net White, Black, and Hispanic populations — that is, total group population minus the number of public housing residents of the same group — and the median contract rent from the U.S. Census, expressed in 2010 dollars.

I apply the decomposition from Equation 8 to map observed changes in rents and group-specific net population into underlying preference parameters. This approach assumes a quasi-linear indirect utility function and that rents adjust endogenously to reflect equilibrium location choices. I then translate these preference parameters into MWTP estimates using Equation 11, which captures the trade-off between neighborhood amenities and housing costs. MWTP values are normalized for interpretability — for example, per 100 additional units, per $1,000 increase in construction cost, or per 10 percentage point change in racial composition — and scaled using average median contract rent by ring, in 2010 dollars.

Finally, I compute the sampling variance of each MWTP estimate using the empirical variance of the underlying reduced-form coefficients. A complete derivation of this decomposition and the variance formula is provided in Appendix A1. Results are presented in Table D5.1.

Table D5.1: MWTP Estimates for post-1970 Public Housing Characteristics

	Panel A: Whites		**Panel B: Blacks**		**Panel C: Hispanics**	
	MWTP (%)	MWTP ($)	MWTP (%)	MWTP ($)	MWTP (%)	MWTP ($)
	PH Tract		PH Tract		PH Tract	
# Apartments	−0.52	−331.78	0.21	131.70	0.13	84.49
	(0.343)		(0.524)		(0.317)	
Construction cost	0.10**	64.07	0.15**	98.01	0.05	32.25
	(0.042)		(0.061)		(0.032)	
# Stories	−0.02	−13.31	−0.12**	−76.63	−0.03	−22.03
	(0.048)		(0.052)		(0.047)	
%Δ Init. Black res.	−0.78***	−499.39	−0.45**	−290.14	−0.26*	−166.26
	(0.207)		(0.229)		(0.157)	
%Δ Init. Hisp. res.	0.24	152.70	−0.14	−87.74	−0.04	−23.43
	(0.204)		(0.259)		(0.137)	
%Δ Init. White res.	0.02	12.89	0.02	13.26	−0.08	−52.27
	(0.174)		(0.274)		(0.120)	
% Ground coverage	−0.66***	−422.13	−0.52*	−336.10	−0.32**	−206.10
	(0.187)		(0.293)		(0.157)	
	1st Ring		1st Ring		1st Ring	
# Apartments	−0.01	−5.70	−0.15	−116.45	−0.23	−177.60
	(0.199)		(0.262)		(0.212)	
Construction cost	0.05*	38.10	0.03	26.22	0.04	29.24
	(0.025)		(0.034)		(0.025)	
# Stories	0.02	12.97	0.04	30.13	0.02	12.45
	(0.021)		(0.035)		(0.031)	
%Δ Init. Black res.	−0.34**	−266.99	−0.12	−89.92	−0.13	−102.42
	(0.156)		(0.116)		(0.082)	
%Δ Init. Hisp. res.	−0.00	−0.47	0.12	91.75	0.07	52.94
	(0.070)		(0.134)		(0.094)	
%Δ Init. White res.	−0.01	−10.57	0.10	74.58	−0.04	−30.57
	(0.086)		(0.153)		(0.069)	
% Ground coverage	−0.42***	−327.23	−0.15	−120.24	−0.15*	−116.66
	(0.099)		(0.145)		(0.089)	

Note: This table reports marginal willingness to pay (MWTP) estimates for public housing characteristics. Panel A presents results for private White residents, Panel B for private Black residents, and Panel C for private Hispanic residents—defined as individuals not living in public housing. Estimates are derived from Equation 10 using coefficients from Figure 6. MWTP values indicate the percentage increase in monthly rent that a household would be willing to pay for a one-unit change in the corresponding attribute. Attribute changes are normalized as follows: 100 units for Apartments, $1,000 for Construction Costs, one storey for Height, and 10 percentage points for Ground Coverage and Racial Composition. The corresponding dollar values are calculated using the average median contract rent (in 2010 dollars) by ring. Estimates are reported separately for public housing tracts (PH tracts) and adjacent areas (1st ring). The sample is restricted to projects constructed before 1970. Standard errors are based on the empirical variance of the underlying population and rent coefficients, as detailed in Appendix A1.
Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.