Urban Renewal and Inequality: Evidence from Chicago's Public Housing Demolitions[*]

Milena Almagro
University of Chicago

Eric Chyn
UT-Austin and NBER

Bryan A. Stuart
Federal Reserve Bank of Philadelphia

December 9, 2022

Abstract

This paper studies one of the largest spatially targeted redevelopment efforts implemented in the United States: public housing demolitions sponsored by the HOPE VI program. Focusing on Chicago, we study welfare and racial disparities in the impacts of demolitions using a structural model that features a rich set of equilibrium responses. Our results indicate that demolitions had notably heterogeneous effects where welfare decreased for low-income minority households and increased for White households. Counterfactual simulations explore how housing policy mitigates negative effects of demolitions and suggest that increased public housing site redevelopment is the most effective policy for reducing racial inequality.

JEL Classification Codes: R23, R28, K42.

Keywords: Urban Renewal, Inequality, Segregation, Endogenous Neighborhood Change.

1 Introduction

Concerns over inequality and the concentration of poverty within cities have prompted governments around the world to invest in community redevelopment and urban renewal programs. These programs invest public resources in specific disadvantaged geographic areas rather than toward low-income individuals directly. In the United States, federal and local governments spend almost $100 billion per year on spatially targeted development programs that aim to revitalize economically distressed communities (Story, 2012; Kline and Moretti, 2014).

Controversy over urban renewal programs often focuses on welfare implications. Policymakers hope that residents of disadvantaged areas benefit from enhanced economic activity and improved local amenities that result from place-based public investments. Yet, critics express concern that revitalization efforts cause increases in the cost of housing that force incumbent low-income residents to relocate to less desirable neighborhoods. Assessing the welfare consequences of urban renewal programs requires understanding both how individuals value neighborhoods and how local housing markets respond to policy.

This paper provides new evidence on the effects of one of the largest spatially targeted redevelopment efforts in the United States: public housing demolitions sponsored by the federal HOPE VI program. The HOPE VI program targeted public housing developments that met standards of extreme physical disrepair, economic distress, and social disorganization. Over a nearly two-decade period, more than $6 billion was spent through the HOPE VI program to transform disadvantaged areas through public housing demolition.

Our approach relies on a structural model of neighborhood demand and supply to quantify the welfare impacts of the HOPE VI program. We focus on the case of Chicago which previously had one of the largest U.S. public housing systems and received substantial HOPE VI funding for building demolition. Between 1995 to 2010, the housing authority in Chicago demolished over 21,000 units of public housing built in neighborhoods throughout the city.

To motivate our model, we highlight stylized facts on how neighborhoods changed after the demolition of public housing in Chicago using U.S. Census data. Between 2000 to 2010, when the vast majority of demolitions occurred, neighborhoods where a larger share of the housing stock was demolished saw substantial increases in the White population share alongside decreases in the share of residents that were Black or Hispanic. Areas with more demolition also saw growth in median household income, median rents, and house values. Redevelopment was also apparent, as the share of newly constructed housing increased more in neighborhoods with more demolitions. When considering the longer-run horizon between 2000 to 2016, there were even larger changes in neighborhood characteristics, suggesting that demolitions had lasting effects. Overall, the evidence indicates that demolitions were followed by migration and broad changes in neighborhoods.

These descriptive findings could be driven by several mechanisms that are key to assessing welfare impacts. For example, one possibility is that individuals explicitly care about the presence of low-income housing in their neighborhood (Diamond et al., 2019). Alternatively, individuals may also care about the demographic characteristics of public housing residents (Bayer et al., 2022). Housing prices could change after demolitions due to either of these channels. Moreover, demolitions could also generate indirect equilibrium effects on prices due to the resorting of individuals and subsequent changes in endogenous amenities.

Our structural approach allows us to quantify the role of these mechanisms and estimate welfare impacts. The model assumes that households have preferences for the demographic and economic characteristics of residents, features of the housing stock, and the presence of public housing. We allow preferences to vary by households' race/ethnicity (non-Hispanic White, Black, Hispanic, and other) and income level (below or above $20,000).[1] The model features several endogenous variables—prices and demographic shares—that allow for a rich set of equilibrium responses to public housing demolitions.

We estimate the model using U.S. Census data that describe the distribution of households across tracts in Cook County, Illinois for the years 2000 and 2010. To identify household preference parameters, we use changes in housing market characteristics and in public housing due to demolitions in distant neighborhoods as instrumental variables in a difference-in-difference framework. This difference-in-difference approach to identification builds on the cross-sectional designs used in Berry et al. (1995) and Bayer et al. (2007). Our strategy leverages the fact that changes in non-adjacent neighborhoods affect prices and demographic shares through substitution patterns. The use of data from 2000 and 2010 allows us to control for time-invariant, unobserved determinants of neighborhood choices that differ across race/ethnicity and income groups. We calibrate the housing supply elasticity using estimates for Chicago from Baum-Snow and Han (2021).

Our estimates of the residential choice model quantify the tradeoffs that households make when deciding where to live. We find that households prefer neighborhoods that have lower rents, a higher share of residents of their own race/ethnicity, and higher income residents. Our estimates are broadly in line with findings from prior studies such as Galiani et al. (2015).[2] All else equal, we also find that households prefer to live in neighborhoods with less public housing. For example, poor White households are willing to pay $139 more in annual rents for a 1 percentage point reduction in the share of public housing in their neighborhood.

---

[1]We use a relatively low threshold given that households located in and near project-based public housing neighborhoods have very low income. These disadvantaged households are plausibly the most affected by demolitions.

[2]Galiani et al. (2015) estimate a neighborhood preference model for low-income, non-White households who participated in the Moving to Opportunity (MTO) housing voucher experiment. They find that households in their sample are willing to pay $122 for a 1 percentage point increase in the share of non-White neighbors. With a related but distinct model of preferences, we find that poor Black households are willing to pay $155 per year to have a 1 percentage point increase in the share of their own race neighbors.

Using the estimated household preferences and calibrated housing supply function, our main analysis examines welfare effects by comparing utility in scenarios with and without HOPE VI sponsored public housing demolitions. Overall, we find that non-poor White households had the most benefits, with their gains from demolition equaling a $230 (2 percent) increase in annual rent equivalent units. In contrast, poor minority households generally saw declines in welfare with negative impacts of -$75 (1 percent) and -$41 (less than 1 percent) for Black and Hispanic households, respectively. Since White households constitute most of the population, there is an overall $127 (1 percent) gain in rent equivalent welfare when we aggregate the welfare of all non-Hispanic White, Black, and Hispanic households. These welfare effects should be interpreted alongside the fact that public housing demolitions affected 5 percent of the neighborhoods in Chicago.

To explore mechanisms, we conduct a partial equilibrium decomposition analysis. Specifically, we create a series of simulations that start with our benchmark counterfactual and selectively allow the features of the model to vary in response to public housing demolitions. In addition to decomposing effects by demographic groups, we analyze welfare separately for renters and homeowners. Our analysis reveals that renters from all demographic groups are worse off from demolitions due to large increases in housing prices that offset welfare gains from the destruction of public housing projects. Welfare losses for White renters are relatively small because they value the equilibrium shifts in demographics that reduced the Black population share in neighborhoods where public housing was demolished. The equilibrium price adjustments are sufficiently large that homeowners for all demographic groups experience welfare increases. When we aggregate impacts across renters and homeowners, the racial disparities in the impact of public housing fully emerge. Both non-poor and poor White households are significantly better off due to their relatively high rates of home ownership. In contrast, poor Black and Hispanic households are worse off.

In addition to studying overall impacts, we use our estimated structural model to quantify spatial spillover effects. While demolitions on average increase rents by 2.4 percent across all census tracts in Cook County, the magnitude of the effects sharply differ based on whether a neighborhood was directly targeted. Neighborhoods with a public housing demolition saw rental prices increase by 13.8 percent on average, while areas without a demolition saw rents go up by 1.9 percent on average. These spillovers, which arise from equilibrium adjustments in the model, generate city-wide impacts even though public housing demolitions occurred in a limited number of neighborhoods.

Finally, a natural question is how the welfare consequences of public housing demolition depend on housing supply responses. To address this, we compare welfare across versions of our model that vary the housing supply elasticity and the extent of additional redevelopment of housing in neighborhoods where demolitions occurred. We find that increasing the housing supply elasticity or the amount of additional redevelopment leads to lower increases in house prices after

demolitions. This in turn raises overall welfare and lowers inequality. Additional redevelopment in neighborhoods where public housing is demolished leads to particularly large effects on prices and inequality because this intervention is more targeted toward neighborhoods where poor and minority households live.[3] Changes in the housing supply elasticity have more muted impacts on prices and correspondingly result in more limited benefits for poor Black households across empirically reasonable values for this parameter.

Overall, this paper makes important contributions to an existing literature that studies the effects of neighborhood renewal policies such as slum clearance programs (Collins and Shester, 2013; Harari and Wong, 2018; Gechter and Tsivanidis, 2020; Blanco, 2021; Blanco and Neri, 2021) and place-based investment programs that aim to revitalize disadvantaged areas (Rossi-Hansberg et al., 2010; Busso et al., 2013). Most directly, our analysis complements prior studies that have estimated short-run neighborhood-level impacts of public housing demolitions in the U.S. (Aliprantis and Hartley, 2015; Sandler, 2017; Tach and Emory, 2017; Blanco, 2021). Our main contribution is that we are the first to study equilibrium impacts of demolitions through the lens of a structural model. As a result, our work provides novel evidence on welfare impacts and explores racial disparities in the effects of place-based housing policies.

Our analysis also contributes to a broad literature that uses structural models of neighborhood preferences to study household sorting and welfare (Bayer et al., 2007; Wong, 2013; Galiani et al., 2015; Bayer et al., 2016; Diamond et al., 2018; Fu and Gregory, 2019; Couture and Handbury, 2020; Davis et al., 2021). A number of studies have estimated impacts of endogenous neighborhood change on welfare, inequality, and segregation (Guerrieri et al., 2013; Diamond, 2016; Almagro and Domınguez-Iino, 2019; Balboni et al., 2020; Couture et al., 2019; Caetano and Maheshri, 2021; Qian and Tan, 2021; Khanna et al., 2022; Su, 2022). Within this body of work, relatively few papers aim to study the effects of housing policies. Most relevant to our analysis are recent studies on the welfare impact of constructing affordable housing through the Low Income Housing Tax Credit (LIHTC) program (Diamond and McQuade, 2019; Davis et al., 2019). We offer two main contributions relative to this prior work. First, the existing evidence suggests that LIHTC is viewed as an amenity in low-income areas and expansions of this form of housing have the potential for welfare gains (Diamond and McQuade, 2019).[4] These findings differ notably from our results which show that relatively large-scale project-based public housing is largely viewed as a disamenity and its removal has the potential to generate substantively important welfare gains. Second, prior work by Diamond and McQuade (2019) and Davis et al. (2019) focuses

---

[3]Our exercise considers a scenario in which the extent of redevelopment is parameterized by the fraction of demolished units that are added to affected neighborhoods. When additional redevelopment exceeds 20 percent, lower-income minority households avoid welfare losses.

[4]Diamond and McQuade (2019) and Baum-Snow and Marion (2009) provide evidence that LIHTC-financed housing developments increase housing prices in low-income areas.

on highly localized impacts of subsidized housing. In contrast, we solve an equilibrium city-wide model and estimate impacts on all neighborhoods. Our innovation is motivated by the large effects of demolitions on demographics and market prices. We find a sizable overall increase in prices and broad spillovers with a majority of the aggregate change attributable to positive impacts in neighborhoods where demolitions did not occur. The spillover results suggest that public housing demolitions may have broadly reshaped the urban landscape in the large U.S. cities (e.g., Baltimore, Atlanta and Philadelphia) that also undertook substantial urban renewal programs during the past three decades.

2   Background

2.1   The Public Housing System in Chicago

Chicago had the third largest public housing system in the U.S. at the beginning of the 1990s. The Chicago Housing Authority (CHA) owned and managed this system, which consisted of high-rise housing developments (also known as "projects") and smaller-scale residential buildings that provided homes specifically for low-income families. High-rise projects consisted of a collection of apartment buildings built in close proximity. Many of these buildings were large structures with approximately 75 to 150 housing units.

Low-income households were eligible to live in public housing if their income was at or below 50 percent of the median income in Chicago. Nearly all residents were Black, and the average household income of public housing residents during this period was $7,000 (Popkin et al., 2000). The resident population was also predominately single-parent, female-headed households.[5]

Public housing was spread across many neighborhoods of Chicago, mostly located in the south and west sides of the city. These areas were predominately Black neighborhoods: the average neighborhood in Chicago with public housing was 70 percent Black in the 1990 Census. Originally, the CHA's buildings had been constructed during the 1950s and 1960s as part of slum clearance and urban development policies pursued by Chicago in the post-World War II era.

By the end of the 1980s, public housing buildings throughout Chicago were in need of serious renovation and repair. Poor conditions in the public housing system stemmed from both the age of the buildings and the fact that funding cuts during the 1980s also complicated building maintenance. More generally, the poor conditions in Chicago's public housing mirrored other major U.S. cities. During the early 1990s, a national commission found that at least 86,000 units of public housing in the U.S. needed major renovation or demolition (U.S. National Commission on

---

[5]According to administrative records, only 6 percent of households living in CHA public housing were headed by a married couple (Popkin et al., 2000).

2.2 Chicago's HOPE VI Demolitions and Redevelopment

City officials in Chicago made plans to demolish public housing as a response to infrastructure problems that had manifested by the 1990s (Popkin et al., 2000).[6] Funding for demolition was provided through the HOPE VI program of the U.S. Department of Housing and Urban Development. Launched in 1992, this program provided support to local city authorities for revitalization and demolition of public housing. Over a nearly two-decade period, the program provided over 400 federal grants to cities across the country.

Chicago was one of the largest recipients of HOPE VI financing. From 1996 to 2003, the city received $83.4 million in grant funding specifically for building demolitions (Aliprantis and Hartley, 2015). Public housing residents were evicted if their building was selected for demolition and received offers for Section 8 housing vouchers that could be used to rent housing from the private market.

As documented in prior studies, Chicago's public housing demolitions took place gradually, with the majority occurring between 1995 and 2010 (Jacob, 2004; Aliprantis and Hartley, 2015; Sandler, 2017; Chyn, 2018; Blanco, 2021). Using administrative records from the CHA, Panel A of Figure 1 plots the total number of public housing units demolished in the 1990s and 2000s. Over 20,000 housing units were demolished during this period, with about 80 percent of demolitions occurring between 2000 and 2010. As shown in Panel B, the timing and intensity of public housing demolition varied widely across the 59 neighborhoods (census tracts) that experienced a public housing demolition during this period.

Figure 2 summarizes the spatial variation in public housing demolitions. In Panel A, we plot deciles of the cumulative number of demolitions in each tract between 1995 and 2010. There is considerable variation: 20 percent of tracts saw fewer than 8 public housing unit demolitions during this period, while 20 percent of tracts saw at least 647 units demolished. The consequences of these demolitions for neighborhoods likely depends on the size of the demolished units relative to the existing housing stock. With this in mind, Panel B focuses on the 59 neighborhoods that experienced a demolition and describes the variation in the intensity of public housing demolition by plotting the total number of units demolished from 1995–2010 as a share of the number of occupied housing units in 1990. The distribution is skewed to the right: most neighborhoods experienced demolitions that account for no more than 20 percent of the 1990 housing stock, but demolitions exceeded 50 percent of the 1990 housing stock in 37 percent of tracts with a

---

[6]Local policymakers believed that they had limited options aside from demolition (Popkin et al., 2000). Few in the city had confidence that the CHA could address housing quality issues due to a series of scandals that revealed housing authorities had mismanaged public funds (Hunt, 2009).

demolition.

What happened to the land after public housing buildings were destroyed? The original plan developed by the CHA was to create "mixed-income" housing in the neighborhoods which formerly featured high-rise public housing (Hunt, 2009). Mixed-income housing would be provided through new construction of both public housing and market-rate units. Yet, as documented in popular media coverage, progress on redevelopment was slow, and the CHA failed to meet its original building goals (Dumke, 2017; Bittle et al., 2017).

Descriptive statistics from land-use data quantify the incomplete nature of redevelopment at former public housing sites.[7] Appendix Table A.1 shows that 38 percent of the lots where public housing was demolished stood vacant and undeveloped in 2010. The land that was redeveloped primarily contained residential housing (40 percent), although some was also occupied by businesses (8 percent) and institutions (4 percent, mostly schools and government buildings).[8] Even by 2015, there was minimal additional progress as the share of vacant land stood at 35 percent.

3   Data

We compile data from two main sources for our analysis. First, we rely on Chicago Housing Authority (CHA) records that provide information on the number of public housing units in a building and the date when the building is destroyed. While we do not observe the date when public housing residents received eviction notices, authorities were required to provide notice at least 5 months in advance. The public housing data contain information on the addresses of both low-rise and high-rise public housing buildings. We map each building address to census tracts and study tract-level measures of the intensity of public housing demolition. Second, we use tabulations from the decennial census and American Community Survey (ACS) to measure residents' race, ethnicity, and median household income, in addition to median rental prices, median home values, and housing unit characteristics. These data cover years 1990, 2000, 2010, and 2016.[9]

Our analysis defines neighborhoods on the basis of census tracts. We create consistently-defined neighborhoods by aggregating census tracts to their 2010 definition, using crosswalks from the Longitudinal Tract Data Base. Our sample includes all tracts in Cook County, Illinois, which contains the city of Chicago. We include tracts throughout Cook County because demolitions in Chicago may have affected neighborhoods in nearby jurisdictions. Our analysis sample, which is

---

[7]We construct these statistics by matching public housing units in CHA administrative data to the Chicago Metropolitan Area for Planning Land Use Inventory in 2010 and 2015 by address.

[8]The remaining categories of land use include industrial (1 percent), roadways and railroads (5 percent), and open space (4 percent).

[9]We use 5-year ACS tabulations, covering 2008–2012 (which we refer to as 2010) and 2014–2018 (which we refer to as 2016). Population counts in 2010 come from the decennial census. Throughout, we use household counts and the race/ethnicity of the head of household.

limited to tracts for which the key variables used in our analysis are not missing, contains 1,240 tracts in Cook County, Illinois.[10] On average, each tract has about 4,000 residents.

4   Motivating Facts

This section documents descriptive facts on the relationship between HOPE VI demolitions and neighborhood outcomes. We document that neighborhoods with a higher extent of public housing demolitions experienced notable shifts in the racial and socioeconomic composition of residents and large increases in housing prices. These descriptive facts motivate our analysis of welfare impacts in Section 7.

4.1   Public Housing Demolitions, Neighborhood Composition, and Housing Prices

We are interested in understanding how residential composition and housing market conditions changed in Chicago's neighborhoods after public housing demolitions. To describe these patterns, we present binned scatter plots of changes in census tract characteristics between 2000 and 2010 against the cumulative number of public housing units demolished during this period as a share of 1990 occupied housing units. We also display changes in neighborhood characteristics from 2000 to 2016 to provide insights on longer-run patterns.

Our analysis begins by examining how the demographic and socioeconomic characteristics of neighborhood residents changed. Each dot in Figure 3 represents the average change in the indicated variable for a given amount of public housing demolition.[11] Panel A shows that the share of residents that are White increased by more in neighborhoods that had a higher intensity of demolition. The slope of the best-fit line for the 2000–2010 change is 0.24, which implies that a 10 percentage point increase in the share of 1990 housing units that were demolished was associated with a 2.4 percentage point increase in the White population share. Correspondingly, we find that minority population shares decreased with the intensity of public housing demolition. Panel B shows that the Black population share fell particularly sharply while Panel C shows that a more muted impact for the Hispanic population share. These demographic shifts were accompanied by changes in log median income. Panel D shows that a 10 percentage point increase in public housing demolition intensity was associated with an 8 percent increase in median household income. Because the vast majority of public housing residents were Black and had low-income,

---

[10]We drop two tracts with public housing from the analysis sample as they are missing data on key outcomes.

[11]Because all tracts with demolitions have a different amount of public housing demolitions as a share of 1990 housing units, this figure shows tract-level values for all tracts with demolitions and a separate average for all tracts with no demolitions.

part of these responses could be explained by the displacement of individuals who were evicted from public housing and moved to a different census tract. That said, the changes in neighborhood characteristics are larger when measured between 2000–2016—a finding that suggests the effects of public housing demolitions are not simply due to mechanical displacement.

Next, Figure 4 studies how the housing market responded to public housing demolitions. Panel A shows that log median rents increased by considerably more from 2000–2010 in neighborhoods where more public housing was demolished. A 10 percentage point increase in the share of 1990 housing units that were demolished was followed by 10 percent faster growth in median rents. Panel B examines changes in median house values given that house prices should better reflect long-run expectations of neighborhood characteristics. The median price of housing is also of interest since it is not affected by the mechanical change in the stock of rented units due to demo-litions.[12] Similar to our analysis of neighborhood demographics and income, the changes in rents and house prices from 2000–2016 are larger than those from 2000–2010. Finally, Panels C and D provide evidence on redevelopment in the neighborhoods which featured public housing demo-litions. Areas with the largest intensity of demolition experienced the largest growth in the share of housing built in the last 10 years and the sharpest declines in the share of housing built more than 30 years ago.[13] We measure all changes in the age of the housing stock between the years 2000–2010 and do not report building age results using the 2016 ACS due to data consistency issues.[14]

Overall, the results in Figures 3 and 4 show that neighborhood demographics, prices, and the housing stock changed notably after public housing demolitions. Specifically, the areas with more demolitions became more White, less affordable, and featured newer housing. In the sections that follow, we develop a structural model to study channels driving these neighborhood changes and quantify the associated consequences for households' welfare throughout Chicago.

5 A Model of Residential Sorting Across Neighborhoods

The demolition of public housing could affect neighborhoods and housing markets in several ways. For example, the attractiveness of a neighborhood to households could depend on the extent of pub-lic housing or the socioeconomic and demographic characteristics of a neighborhood's residents.

---

[12]Public housing is rented, and so demolitions could mechanically increase the median rent by eliminating housing units at the bottom of the distribution. Home prices are not affected by this issue.

[13]Appendix Figure A.1 reports additional results for the share of housing units built 11–20 and 21–30 years ago. There are more muted impacts for these housing stock outcomes because the oldest housing units (which included public housing) were more likely to be redeveloped.

[14]We use ACS data from 2014–2018 for the 2016 numbers, and we are only able to see the share of housing units built in 2014 or later, 2010–2013, 2000–2009, 1990–1999, and so on. Unfortunately, these bins do not line up with the 10-year bins available for the 2000 and 2010 Census.

If this is the case, then households might make different location choices after demolitions, leading to changes in equilibrium housing prices and endogenous amenities. To study the channels driving changes in neighborhoods after demolitions and assess welfare consequences, this section develops a model of equilibrium sorting by combining a discrete choice model of residential demand (Bayer et al., 2007) with a model of housing supply.

5.1 A Model of Neighborhood Demand and Housing Supply

Households of race-by-income group k choose their neighborhood location at time t by solving the following problem:

$$\max_j V_{ijt}^k = \delta_{jt}^k + \epsilon_{ijt}^k,$$

where δ_{jt}^k is the component of indirect utility for neighborhood j that is common to all households of group k, and ϵ_{ijt}^k is an idiosyncratic shock that is assumed to be an i.i.d. type I Extreme Value. The common component of indirect utility is:

$$\delta_{jt}^k = \alpha_p^k \ln(p_{jt}) + \alpha_b^k b_{jt} + \alpha_h^k h_{jt} + \alpha_{Inc}^k \ln(Inc_{jt}) + \alpha_{PH}^k PH_{jt} + \theta^k x_{jt} + \xi_{jt}^k, \tag{1}$$

where p_{jt} is the rental price of housing, b_{jt} and h_{jt} are the share of households that are Black or Hispanic, Inc_{jt} is median household income, PH_{jt} is public housing as a share of housing stock in tract j, x_{jt} is a vector of exogenous observable neighborhood characteristics such as features of the housing stock or land-use shares across several categories, and ξ_{jt}^k is a scalar that summarizes unobservable neighborhood characteristics. Preference parameters, $\alpha^k \equiv (\alpha_p^k, \alpha_b^k, \alpha_h^k, \alpha_{Inc}^k, \alpha_{PH}^k, \alpha_x^k)$, as well as neighborhood unobserved quality, ξ_{jt}^k, may differ arbitrarily across groups. We use vectors (e.g., $\mathbf{p}$, $\mathbf{b}$, and $\mathbf{h}$) to represent aggregates across the set of J-many neighborhoods (i.e., $\mathbf{p_t} \equiv (p_{1,t}, \ldots, p_{J,t})$). We assume that home prices are equal to the present discounted value of rents, and therefore homeowners face the same optimization problem as renters.[15]

Given our distributional assumption on ϵ_{ijt}^k, the probability that a group-k household chooses to live in neighborhood j is:

$$\mathcal{P}_{jt}^k(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_{\mathbf{t}}^{\mathbf{k}}; \alpha^k) = \frac{\exp\left(\delta_{jt}^k\right)}{\sum_{j'} \exp\left(\delta_{j't}^k\right)}, \tag{2}$$

where we include log median household income and the public housing share in $\mathbf{x_t}$ to conserve on notation in the above expression. The demand for living in neighborhood j equals the total number

---

[15]Our assumption is in line with prior studies that similarly do not separately model the decision to buy or rent a home (Bayer et al., 2016; Diamond and McQuade, 2019).

of households, across all groups, that want to live in j:

$$\mathcal{D}_{jt}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha) = \sum_k \mathcal{P}_{jt}^k(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi^\mathbf{k}{}_t; \alpha^k) N_t^k, \tag{3}$$

where N_t^k is the total number of group k households in Illinois, which we take as exogenous.

We close the model by assuming a housing supply curve. Our approach follows Davis et al. (2019) and assumes a supply relationship based on estimates of housing elasticities from the literature. Specifically, we assume that the number of housing units supplied in neighborhood j is an isoelastic function of the price:

$$\mathcal{S}_{jt}(p_{jt}) = \theta_{jt} p_{jt}^\psi, \tag{4}$$

where θ_{jt} is a supply shifter and ψ is the supply elasticity which we obtain based on prior studies (as detailed below).

An equilibrium of this model occurs when prices and demographic characteristics of neighborhoods lead to market clearing. More formally, the equilibrium prices $\mathbf{p_t^*}$ and demographic shares $(\mathbf{b_t^*}, \mathbf{h_t^*})$ are vectors that satisfy the fixed-point defined by the following system of equations:

$$\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha) = \mathcal{S}_{jt}(p_{jt}^*) \qquad\qquad \forall j = 1, ..., J \tag{5}$$

$$\frac{\mathcal{D}_{jt}^B(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha)}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha)} = b_{jt}^* \qquad\qquad \forall j = 1, ..., J \tag{6}$$

$$\frac{\mathcal{D}_{jt}^H(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha)}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha)} = h_{jt}^* \qquad\qquad \forall j = 1, ..., J, \tag{7}$$

where $\mathcal{D}_{jt}^B(\cdot)$ and $\mathcal{D}_{jt}^H(\cdot)$ are the equilibrium number of Black and Hispanic households in neighborhood j. The existence of endogenous variables besides housing prices arises from the fact that a neighborhood's demographic characteristics are the result of household location decisions. As we show below, this richer equilibrium concept is important for understanding the effects of demolitions on neighborhoods.[16]

6 Quantification of the Model

6.1 Approach

To study the consequences of public housing demolitions using our model, a necessary step is to obtain estimates of the household preference parameters. As is standard, the indirect utility of the model's outside option—living outside of Cook County in Illinois—can be normalized to be equal

---

[16]Appendix B describes the numerical procedure for solving for the equilibrium with endogenous amenities.

to zero: $\delta_{0t}^k = 0$.[17] Equation (2) then implies the following relationship that can be taken directly to aggregate data (McFadden, 1974; Berry, 1994):

$$\log\left(\frac{\mathcal{P}_{jt}^k}{\mathcal{P}_{0t}^k}\right) = \alpha_p^k \ln(p_{jt}) + \alpha_b^k b_{jt} + \alpha_h^k h_{jt} + \alpha_{Inc}^k \ln(Inc_{jt}) + \alpha_{PH}^k PH_{jt} + \theta^k x_{jt} + \xi_{jt}^k. \quad (8)$$

To measure the choice probabilities, we rely on census and ACS data on household counts for each group k for each tract in Cook County. We focus on eight race-by-income groups defined by dividing each of four major categories of race/ethnicity (non-Hispanic White, Black, Hispanic, and other) into poor and non-poor households (those with income below and above \$20,000). We focus on this definition for poor households given that the households that typically lived in and around the public housing neighborhoods had very low household income.[18,19]

Notably, we face an empirical challenge for obtaining estimates of the choice probabilities for each of these eight groups. Specifically, these probabilities could be estimated using the share of households of group k that reside in each tract:

$$\hat{P}_{jt}^k \equiv \frac{\text{\# Residents of group } k \text{ in tract } j \text{ at time } t}{\text{\# Residents of group } k \text{ in Illinois at time } t}. \quad (9)$$

In practice, there are two concerns regarding the shares in equation (9). First, the tract-level population measures are subject to measurement error due to the fact the census and ACS data is based on a small subsample of the U.S. population.[20] Second, the share estimates for a given demographic and racial group may take values of 0 or 1 which is inconsistent with the assumed logit errors in our structural model. As a result, we smooth choice probabilities by taking a weighted average of our frequency estimates across census tracts:

$$\tilde{P}_{jt}^k = \sum_n w_{jn} \hat{P}_{nt}^k.$$

As in Scott (2013), the weight is inversely related to the distance between the centroids of tracts and normalized so that weights add up to one:

$$w_{jn} = \left(\frac{1}{1 + \text{dist}(j, n)}\right) \Big/ \left(\sum_{j'} \frac{1}{1 + \text{dist}(j', n)}\right).$$

---

[17]Given this outside option, we have an open-city model. In this way, the model features endogenous population flows in and out of Cook County. We assume that the population of Illinois is exogenous and determined outside our model.

[18]Neighborhoods subject to demolition in Chicago had an average poverty rate of 54% in 2000 (Aliprantis and Hartley, 2015). The poverty line in 2000 for a two-adult household with two children was just under \$20,000 in 2000.

[19]We do not focus on more granular definitions of demographic groups since we face a trade-off between having more heterogeneity in preference parameters and less precise estimates of choice probabilities when groups have fewer individuals.

[20]The 2000 Census collected income information for about 17 percent of all households, while the combined 2008–2012 ACS data contain about 5 percent of households (1 percent in each year).

For the independent variables of the model, we measure p_{jt} using median gross rents (equal to contract rent plus the cost of utilities) from the census and ACS data. The shares of households that are headed by Black or Hispanic individuals, b_{jt} and h_{jt}, as well as median household income, Inc_{jt}, also come from census and ACS data. We measure the share of housing units that are public housing using public housing estimates from the CHA and total housing estimates from the census and ACS.[21] The vector x_{jt} includes several variables that could influence the attractiveness of a neighborhood to residents: the share of housing units that are owner-occupied, the log median number of rooms in housing units, the log median year that housing units were built, as well as the share of land allocated to various uses (residential, construction, industrial, other urban, infrastructure, agriculture, open, and water). The land use variables help us control for the industrial composition of different areas and access to job opportunities.

For estimation, our approach addresses two main threats for credible identification of the preference parameters of our model of residential choice. First, we include a series of fixed effect terms by estimating our model using repeated cross sections from 2000 and 2010. Specifically, we include tract fixed effects, λ_j^k, that account for fixed characteristics that do not change over time such as distance to the central business district (Nevo, 2001). We also control for common shocks to all tracts by including year t fixed effects, λ_t^k. Both fixed effects vary arbitrarily by race-and-income group. The empirical fixed effect specification we use to estimate household demand parameters is:

$$
\log\left(\frac{\tilde{P}_{jt}^k}{\tilde{P}_{0t}^k}\right) = \alpha_p^k \ln(p_{jt}) + \alpha_b^k b_{jt} + \alpha_h^k h_{jt} + \alpha_{Inc}^k \ln(Inc_{jt})
$$

$$
+ \alpha_{PH}^k PH_{jt} + \theta^k x_{jt} + \lambda_j^k + \lambda_t^k + \tilde{\xi}_{jt}^k. \tag{10}
$$

Second, an additional concern when estimating equation (10) is that some observable characteristics may be correlated with changes in unobserved neighborhood quality, $\tilde{\xi}_{jt}^k$. For example, some neighborhoods might become relatively more attractive over time, leading to greater house price appreciation, and ignoring this confounder would lead to an upward bias in the coefficient for prices. Besides housing prices, demographic shares and median household income are equilibrium outcomes that may depend on $\tilde{\xi}_{jt}^k$. Therefore, we also expect OLS estimates of these variables to be biased.

To overcome this second identification challenge, we construct instruments using measures of public housing shares and housing characteristics in neighborhoods that are farther than 3 miles

---

[21]Data from the CHA provide us with information about the number of public housing units demolished in each period. The information on the stock of public housing units in each period appears to be less reliable. However, the stock of units in year $t \in \{2000, 2010\}$ is equal to the stock of units in year 1990 minus the total number of demolitions between 1990 and year t. Because the stock of units in year 1990 does not change over time, it is absorbed by the fixed effects included in our regressions below. As a result, we are able to include in our regression a variable that is equivalent to the share of housing units that are public housing in each period.

13

away. This approach broadly follows Bayer et al. (2007). In our particular application, the instrument vector z_{jt} contains separate averages of the public housing share, median number of rooms, and median year built variables for neighborhoods that are 3–5, 5–10, and 10–20 miles away. The relevance condition is satisfied because neighborhoods are substitutes for each other. As substitutes, shifts in the characteristics of other neighborhoods alter equilibrium prices and residential composition across neighborhoods. For example, consider the high-income neighborhood of Lincoln Park in Chicago. New construction in this area may attract households who prefer newer buildings and reduce demand for other relatively high-income neighborhoods such as Hyde Park that may be located further away. In general, we expect that new construction or other changes in the housing stock may change demand due to substitution and thereby affect housing prices in areas that did not directly experience new construction. Moreover, if different demographic groups vary in their valuations for different housing characteristics, we similarly expect the same mechanism to act as a shifter of the demographic composition of areas that did not experience new construction.

More formally, the exclusion restriction is satisfied if changes in the public housing share and physical housing characteristics in distant neighborhoods are uncorrelated with unobservable neighborhood trends:

$$\mathbb{E}[\tilde{\xi}_{jt}^k z_{jt} | PH_{jt}, x_{jt}, \lambda_t^k, \lambda_j^k] = 0.$$

The plausibility of this exclusion restriction is motivated by three distinct considerations. First, conditional on the extent of public housing that existed around 2000, changes in public housing from 2000 to 2010 were driven in part by idiosyncratic factors like pipes bursting which led housing authorities to decide to demolish a particular building. Second, changes in the housing stock in distant neighborhoods (e.g., due to redevelopment) were unlikely to depend on unobservable trends in a given neighborhood. Third, consistent with these previous points, Figure 5 shows that changes in neighborhood rents from 1990–2000 are uncorrelated with both the extent of public housing demolitions (Panel A) and our IV-based predictions for the changes in neighborhood rents (Panel B) from 2000–2010. Similarly, there is also no significant relationship between these sources of identifying variation and changes in house values (Panels C and D).

To increase the first stage power, we use a three-step approach following previous studies (Bayer et al., 2007; Davis et al., 2019). First, we estimate preference parameters using equation (10) and the already-discussed instruments. Second, we solve for equilibrium rents (p_{jt}) and location choices ($\mathcal{P}_{jt}^k$) under the assumption that there are no unobserved time-varying determinants of neighborhood quality (i.e., $\tilde{\xi}_{jt}^k = 0$). Using the location choices, we construct the share of each neighborhood that is composed of each race-by-income group k. These simulated instruments exploit the equilibrium conditions of the model to concentrate the exogenous determinants of rents and demographic shares into a functional form that is particularly predictive of actual rents and de-

mographic shares. Finally, we estimate equation (10) using as instruments the simulated log price and the share of households in each race-by-income group. The three-step approach increases the first stage Kleibergen-Paap F-statistic from 1.1 to 10.7.

Finally, in addition to estimating household preferences, our analysis requires us to take a stand on housing supply responses. We calibrate the housing elasticity ψ in equation (4) using estimates from Baum-Snow and Han (2021), who estimate tract-level supply elasticities for Chicago between 0.106 and 0.220. In our baseline analysis we take the middle point within that range and set $\psi = 0.163$. In our counterfactual exercises in Section 9, we explore a range of other values and show how welfare results change with respect to the calibrated elasticity. Finally, to calibrate the intercept of the supply curve, θ_{jt}, we combine the supply curve in equation (4) with equilibrium quantities under the simplifying assumption that the unobserved demand component $\tilde{\xi}_{jt} = 0$:[22]

$$\hat{\theta}_{jt} = \mathcal{D}_{jt}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \mathbf{0}; \hat{\alpha})/p_{jt}^{\psi}.$$

6.2   Household Preference Results

Table 1 presents instrumental variable estimates of equation (10).[23] Panel A presents results for poor households (those with income below $20,000), while Panel B shows results for non-poor households (with income of $20,000 or more). As expected, we find that all households dislike paying more for housing. We also estimate preferences that are consistent with racial homophily: coefficients on the Black and Hispanic population shares are negative for White residents and positive for Black and Hispanic residents. Conditional on a neighborhood's cost of living and demographic composition, households prefer to live in neighborhoods where the median household income is higher. Finally, we also find that the presence of public housing is a disamenity. Notably, a comparison of these IV-based results and OLS results in Appendix Table A.2 suggests that failure to account for the potential endogeneity of prices and demographic shares leads to upward bias in price coefficients and considerably larger estimated willingness to pay for demographic characteristics.[24] This upward-bias is consistent with prices being positively correlated with unobservable demand shocks that are addressed in our IV approach.

There are two main caveats for the interpretation of the estimated coefficients in Table 1. First, we interpret these results as reduced-form parameters that may reflect the combined impact of

---

[22]The correlation between observed equilibrium quantities from the census and ACS data and our implied equilibrium quantities is 0.98. Our welfare results remain virtually unchanged if we use observed equilibrium quantities to calibrate θ_{jt}.

[23]We present the main results for non-Hispanic White, Black and Hispanic households for clarity. Appendix Table A.4 reports results for non-poor and poor other race/ethnicity households. The other race/ethnicity group constitutes just 7 percent of the households in Cook County.

[24]For example, the willingness to pay for an increase in the share of Black neighbors based on the OLS estimates is an order of magnitude larger than the corresponding value from the IV-based estimates for several groups.

additional preferences that we do not explicitly model. For example, White households might prefer to live in neighborhoods with a higher White population share because of racial animus, preferences for public goods that are associated with demographic composition, or preferences for particular types of consumption amenities (Almagro and Domınguez-Iino, 2019). Similarly, our estimates for demographics or public housing may partly reflect preferences and beliefs about the presence of crime.[25] In choosing the number of arguments to include in the indirect utility function, we balance the trade-off between adding potentially relevant variables and retaining a parsimonious model whose estimates are readily interpretable.

Second, while we refer to the estimates as reflecting "preferences," they also could reflect constraints. For example, poor households might be more sensitive to housing prices not because of inherent differences in what they value, but simply because they are more financially constrained. In a similar vein, the location choices of Black and Hispanic households could be constrained by discrimination in the housing market (e.g., Christensen and Timmins, 2021).

To compare preferences across groups, we calculate the implied willingness to pay for different neighborhood characteristics.[26] Poor White residents are willing to increase their annual rent by about $51 and $41 for 1 percentage point *decreases* in the share of residents that are Black or Hispanic, respectively. Conversely, poor Black and Hispanic residents are willing to increase their rents by $155 and $176 respectively for 1 percentage point *increases* in the share of their own demographic group. Non-poor households are willing to pay even larger amounts to live in a neighborhood with the same demographic group. All race-by-income groups have positive willingness to pay for reductions in the public housing share.

How do these estimates compare to previous studies? One comparison for our analysis comes from Galiani et al. (2015) which used data from the MTO housing voucher experiment to estimate a similar model of neighborhood preferences for households living in public housing. They focus on non-White households and estimate an average annual willingness to pay of $122 for a 1 percentage point increase in the share of non-White neighbors. This finding is consistent with preferences for neighbors of the same race and quantitatively similar to the willingness to pay for same-race neighbors for poor Black households.

Appendix Tables A.3–A.6 show that the preference parameter estimates are similar across a range of alternative specifications. First, we explore potential sensitivity to spatial spillovers by adding variables that measure neighborhood conditions (log median number of rooms, log median

---

[25]Given that we do not explicitly model crime, one may view our parameterized model as a first-order approximation where the estimates are a function of underlying preferences over demographics, crime, as well as the parameters underlying the crime production function. We refer interested readers to Khanna et al. (2022) for a spatial framework that incorporates a crime production function.

[26]It is more difficult to interpret cross-group differences in the coefficients in Table 1 because the coefficients are only identified relative to the variance of the idiosyncratic Type I extreme value error term, which can vary across groups.

year built, and public housing share) in neighboring tracts that are less than 1 mile away, 1–2 miles away, and 2–3 miles away (column 2). This test is motivated by the idea that similarity of these results would suggest that there is negligible omitted variation in spatial spillovers that threatens identification in our main specification. Second, we estimate regressions that control for a measure of crime least subject to measurement error, the tract-level homicide rate (column 3). Robustness to this specification would imply that our reduced-form utility flow parametrized as a function of the demographic composition is a reasonable first-order approximation that reflects how changes in crime rates may drive residential choice. Third, to demonstrate the robustness in our IV approach, we show results where we vary the definition of the "further away" neighborhoods used to construct our instruments (columns 4 and 5).[27] Fourth, we also explore alternative specifications which move away from the common trends assumption implied by the inclusion of λ_t^k in our preferred model. Specifically, we augment our main specification by including interactions between group-specific fixed effects for the year and measures of 1990 neighborhood characteristics as well as interactions between group-specific fixed effects for year and the 1990–2000 changes in the same neighborhood characteristics. Fifth, we also estimate a model on a sample that excludes all census tracts within one mile of the Cabrini-Green public housing project (column 7) due to concern that unobserved trends in gentrification may confound our ability to obtain unbiased estimates. The results from Appendix Tables A.3–A.6 show that the estimated preference parameters are quite robust.

Finally, we also explore heterogeneity in preferences over public housing to further characterize the interpretation of our main estimates. Our analysis is motivated by findings from Diamond and McQuade (2019) which show that affordable housing constructed by the Low-Income Housing Tax Credit (LIHTC) program is viewed as a disamenity only in high-income neighborhoods. Appendix Table A.7 provides results based on a model that allows the preference over public housing to vary with the income level of the neighborhood. In particular, we divide tracts in Cook County into deciles based on 1990 median household income, and we modify our baseline specification to allow the public housing coefficients to differ for the bottom (first) and remaining deciles (second through tenth). The results show that public housing is viewed as a disamenity in all neighborhoods, although White households view public housing as a larger disamenity in poor neighborhoods.

---

[27]Our main specification focuses on public housing and housing characteristics in the neighborhoods that are 3–5, 5–10, and 10–20 miles away. Columns 4 and 5 show results where the instruments are based on the relatively closer neighborhoods that are 2–3 and 3–5 miles away or 2–3, 3–5, and 5–10 miles away, respectively.

7 Welfare Impacts of Public Housing Demolitions and Revitalization

7.1 Framework

What impact did public housing demolition have on welfare? To answer this question, we combine the estimated household preferences and the calibrated housing supply elasticity with our model from Section 5. Using our estimated demand parameters and our calibrated supply curve, we calculate static equilibria and welfare in several counterfactual scenarios.[28] Our goal is to compare welfare from counterfactual scenarios to a baseline scenario where welfare is calculated using observed outcomes from the 2010 Census. This baseline scenario corresponds to the actual situation where Chicago destroyed public housing through the HOPE VI program.

Throughout the text, we focus on household welfare unless otherwise specified. As noted above in our discussion of our model of neighborhood demand, we assume that renters and homeowners within a given group k have the same preferences and home prices are equal to the present discounted value of rents. Hence, homeowners of group k make the same location choices as their counterpart renters. While renters and homeowners receive the same utility flow from a given neighborhood, an important additional consideration is that homeowners' welfare also depends on the flow of rental income from their housing portfolio. In Chicago, this is a realistic concern given that a large portion of the housing stock is owner-occupied. Therefore, we compute household welfare as the sum of two components: the consumer surplus enjoyed by renters and homeowners and the rental income of homeowners.

Using the estimated preference parameters, a specified set of neighborhood characteristics $(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x})$, and the properties of the Type I Extreme Value distribution of the idiosyncratic shock, we compute the average renter consumer surplus for households of group k in closed-form solution as follows:

$$\mathcal{CS}^k(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi^{\mathbf{k}}; \alpha^k) = \log \left(\sum_j \exp \left(v_{jt}^k(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi^{\mathbf{k}}; \alpha^k) \right) \right),$$

where $v_j^k(\cdot)$ is indirect utility specified in equation (1). In all of our equilibrium analysis, we assume that the unobservable demand factor is equal to its conditional mean across all groups, that is, $\tilde{\xi}_{jt}^k = 0$ for all j, t and k. Therefore, to simplify notation, we suppress that term inside our

---

[28]Our static framework is suited to capture long-run outcomes, where the economy can be thought to be in steady-state. While dynamic considerations and short-run effects could be important, studying these issues would require data that identify individuals' race, socioeconomic status, and location choices over time. The cross-sectional nature of the census and ACS data does not allow us to construct a longitudinal panel of location choices. Other datasets (such as those from Infutor) contain limited information about individuals' race and income level and may not have adequate representation of lower-income households.

consumer surplus measure.[29]

To compute renter welfare changes from a counterfactual world $(\mathbf{p}^1, \mathbf{b}^1, \mathbf{h}^1, \mathbf{x}^1)$ relative to a baseline scenario $(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^0)$ in monetary terms, we rely on the notion of a rent equivalent. We define the group-specific rent equivalent, RE^k, as the increase in rent that is necessary to leave the household indifferent with respect to the baseline values as follows:

$$\mathcal{CS}^k(\mathbf{p}^1 + RE^k, \mathbf{b}^1, \mathbf{h}^1, \mathbf{x}^1; \alpha^k) = \mathcal{CS}^k(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^0; \alpha^k),$$

where positive values of the rent equivalent are associated with higher welfare in the counterfactual world.

Focusing on the previous rent equivalent measure is useful because it allows us to measure changes in renter consumer surplus in monetary terms. Therefore, we can readily measure the welfare change for homeowners as the sum of the the rent equivalent and the change in rents, which accrues to homeowners as rental income.

Finally, we assume that all homeowners own a fully diversified portfolio of housing for simplicity, so that all homeowners receive the same increase in rental income. Under this assumption, the overall welfare effects of group k is defined as the weighted average of welfare changes across homeowners and renters:

$$RE^k + s_{home}^k \cdot \Delta \bar{r},$$

where s_{home}^k is the percentage of households in group k who are homeowners and $\Delta \bar{r}$ is the average rent change across tracts in Cook County.

7.2   Assessing Model Fit

Before describing the welfare consequences of public housing demolitions, we conduct two validation exercises to assess how well our model fits equilibrium rental prices. In both exercises, we focus on the explanatory power of the explicitly-modeled elements by setting the unobserved time-varying index of neighborhood quality equal to zero (i.e., $\tilde{\xi}_{jt}^k = 0$ for all j, t, and k). Rental prices are a particularly useful outcome because they depend on both the demand and supply components of the model.

Our first analysis in Figure 6 plots actual log rents in census tracts in 2000 or 2010 against log rents that are implied by the associated model equilibrium (with $\tilde{\xi}_{jt}^k = 0$). In Panels A and B, the intercept of the housing supply curve is estimated using the number of housing units implied by the demand system. As a result, actual and simulated rents in these panels differ only because of

---

[29]Alternatively, we could incorporate the estimate $\hat{\tilde{\xi}}_{j,t}^k$ for $t = 2000, 2010$ into the utility function. However, this would implicitly require the realization of this component to remain unchanged across different scenarios, which is arguably a stronger assumption. That said, it is worth noting that results are qualitatively and quantitatively similar when we set the unobservable demand factors equal to the residuals of equation 10.

the time-varying unobserved demand factor, $\tilde{\xi}_{jt}^k$. The actual and simulated data are nearly identical, which implies that the explicitly included variables in the simulation (i.e., everything aside from $\tilde{\xi}_{jt}^k$) explain nearly all of the relevant variation in equilibrium prices. In Panels C and D, the intercept of the housing supply curve is instead estimated using the observed number of housing units in the census/ACS data (smoothed across tracts, to be consistent with the smoothed choice probabilities). As a result, differences between the actual and simulated data can arise because of the unobserved demand factor, $\tilde{\xi}_{jt}^k$, and general model misspecification. Using the observed supply, the simulated data explain 97 percent of the variation in log rents in 2000 and 96 percent of the variation in 2010. These results demonstrate a high degree of in-sample model fit.

Second, we conduct a more-stringent, out-of-sample exercise that examines whether the estimated model—which is based on 2000 and 2010 data—can accurately predict rents in 1990. For this analysis, we use the coefficients and tract fixed effects estimated using 2000–2010 data and the exogenous observed neighborhood characteristics in 1990. In addition, we assume that the housing supply shifter, θ_{jt}, is the same in 1990 and 2000. The equilibrium definition in equations (5)–(7) allows us to solve for the endogenous variables in this exercise. Importantly, we do not use any data from 1990 on the endogenous variables in this procedure. Our test is a comparison of the resulting equilibrium rents simulated out-of-sample for 1990 against the actual rents. The results in Figure 7 show that there is an almost one-to-one relationship between actual and simulated rents on average. Moreover, the simulated rents explain 70 percent of the cross-tract variation in actual rents. This out-of-sample validation exercise underscores the strong fit of the model.

7.3 Main Results

We begin by calculating the average change in each group's welfare due to public housing demolitions. To do this, we compare welfare under the actual state of the world in 2010 (where demolitions occurred) to a counterfactual version of 2010 in which the public housing share in each tract is held constant at its level in 2000.[30] We solve for the equilibrium rents, demographic shares, and households' location choices in this year 2010 counterfactual.[31]

Figure 8 reports the rent equivalent changes in welfare per household due to public housing

---

[30]To construct the counterfactual with no public housing demolitions, we remove residents of public housing from the market. To do so, we reduce the number of poor Black households in 2010 by the number of occupied public housing units that were demolished (approximately 15,000). This allows us to capture the fact that the public housing demolitions that had been completed as of 2010 had increased the number of households that demand market-rate housing in the observed world. Because the number of residents displaced from public housing is small relative to the total number of residents in Chicago, results are very similar when our analysis does not decrease the number of households in the counterfactual. The correlation between the counterfactual change in rents due to public housing demolitions under these alternative assumptions on the number of households in the private market is 0.99, with a maximum absolute deviation of 3.6 percent.

[31]To compare the observed 2010 outcomes with 2010 counterfactual outcomes, we assume that the outside option (i.e., a location in Illinois but outside Cook county) remains unchanged.

demolitions for each group.[32] The rent equivalent for non-poor White households is the highest among all groups and implies that these households see an increase in utility due to demolitions that would be offset only by a $230 increase in their annual rents (representing a 2 percent increase relative to the mean annual rent in Cook County in 2010). Poor White households also experience an increase in utility due to public housing demolitions but by a smaller amount at $113. In contrast, poor Black and Hispanic households are worse off because of demolitions. The decrease in utility is equivalent to -$75 per year for poor Black households and -$41 for poor Hispanic households. To understand the overall welfare effects, we combine these group specific impacts into a 2010-population weighted average. Overall, we estimate that public housing demolitions increased the average welfare of non-Hispanic White, Black, and Hispanic households by $127 (1 percent) due to the fact that non-poor White households are the largest demographic group in our context.

7.4   Addressing Multiplicity of Equilibria

A potential concern for the interpretation of our main welfare results is that our model may feature multiple equilibria. Given that we treat neighborhood demographic variables as endogenous, the model implicitly features agglomeration forces. In general, if congestion forces are dominated by agglomeration forces, the model may exhibit multiple equilibria (Bayer and Timmins, 2005). The presence of multiple equilibria thus depends on preference parameter estimates.

We explore the presence of multiple equilibria in two ways. First, we focus on solving for the equilibrium for Cook County in 2010 and initialize our equilibrium solver from 1,000 different starting values. We find only negligible differences across the fixed point that defines the equilibrium conditions. Second, we follow Bayer and Timmins (2005) and initialize our equilibrium solver by setting demographic shares in different neighborhoods at extreme values. With this alternative approach we also find the same solution for the fixed point in our equilibrium definition. Overall, we take these heuristic results as suggestive evidence that the model does not feature multiple equilibria with our estimated preference parameters.

Although we do not provide a formal proof to rule out multiple equilibria in our context, it is possible that certain combinations of model primitives, such as preference parameters or exogenous neighborhood characteristics, could lead to a unique equilibrium. Concretely, the results from Bayer and Timmins (2005) show that a unique equilibrium is more likely when the choice set is larger, the locations are ex-ante more distinct, or the model features groups with strongly differing preferences. In our case, our choice set contains more than 1,200 locations, there is large variation in the characteristics of locations, and notably different willingness to pay across demographic

---

[32]We also estimate the welfare effects for other race/ethnicity households. The change in rent equivalent welfare is -$221 and $30 for poor and non-poor other race/ethnicity households, respectively.

groups.

7.5 Decomposition Analysis

To better understand the overall impact of public housing demolitions on welfare, Table 2 provides decomposition results that are based on estimating several additional counterfactual scenarios. Each row reports rent equivalent statistics based on comparing a scenario in which there are no public housing demolitions to a scenario that selectively varies which endogenous features of the model are permitted to adjust when demolitions occur.[33] The counterfactuals are defined to selectively highlight the quantitative importance of various equilibrium channels that are embedded in our model. Note that scenarios which allow only some endogenous features to adjust while others remain constant are partial equilibrium results.[34]

Panel A focuses on welfare for renters alone. The first row reports results where the counterfactual considered is one in which public housing is destroyed but demographics and rents are fixed at the the counterfactual 2010 levels with public housing. Because all groups view public housing as a disamenity, the rent equivalent numbers from destroying public housing in the first row are positive. The results in the second and third rows show that changes in neighborhood demographics notably contribute to the effects of public housing demolitions. In the second row, we consider a scenario in which neighborhood composition changes only because of the removal of public housing residents.[35] This mechanical change in neighborhood composition increases the utility of White households, who prefer to live in neighborhoods with fewer Black residents, and decreases the utility of Black and Hispanic households, who value living near Black neighbors. The third row allows for broader changes in neighborhood demographics by allowing all households to re-optimize their location choice in response to public housing demolitions. This resorting leads to higher utility for Black and Hispanic households while also resulting in lower utility for White households. These results show that demolitions disrupted areas that had a favorable demographic composition for minorities, but equilibrium resorting allows minority groups to partially recreate the demographic landscape of the disrupted communities. Finally, the fourth row illustrates the importance of price adjustments for renters. Households from all groups are worse off in

---

[33]For example, the rent equivalent shown in the first row of Table 2 is computed as:

$$\mathcal{CS}^k(\mathbf{p}^0 + RE^k, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^{PH,own}; \alpha^k) = \mathcal{CS}^k(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^0; \alpha^k)$$

where $\mathbf{x}^{PH,own}$ differs from $\mathbf{x}^0$ only by incorporating public housing demolitions in each tract. Diamond (2016) performs a similar gradual decomposition exercise.

[34]The "all channels" results in rows 4 and 5, as well as the results in Panel C, are based on a general equilibrium analysis.

[35]We assume that all public housing residents are Black, which is approximately true in the context of Chicago during our study period (Popkin et al., 2000; Chyn, 2018).

this scenario as our simulations find that demolitions increased rents substantially.

We summarize the effects on homeowners in Panel B. As noted in Section 7.1, renters and homeowners have the same preferences, but the welfare of the latter is a function of rental income. In line with our relatively large estimated impact of demolitions on rental prices, the results show that homeowners from all race and income groups have improved welfare outcomes.

Finally, the remaining rows of Table 2 summarize both the average and aggregate population impacts of public housing demolitions by group. The average results display the main welfare results from the general equilibrium exercise in which all endogenous channels operate and we consider both renters and homeowners. Comparing this row to the previous intermediate scenarios shows that price adjustments and homeownership rates notably shape racial disparities in the effects of demolitions. Intuitively, the pattern of results stems from the fact that homeownership rates vary substantially across demographic groups. For example, in 2010, the year of our counterfactual, homeownership rates vary from 81 percent for non-poor White households to 19 percent for poor Black households.[36] Due to this disparity in homeownership, the rent equivalent welfare gain of non-poor White households is no longer negative ($230 versus -$21) and the gap between this group and poor Black households increases by 170%. Overall, when we scale up to the population for each group, we find sizable gains for poor and non-poor White households of $13.6 and $180.9 million, respectively. Losses for low-income Black and Hispanic households total $13 million.

8   Impacts on Neighborhoods Throughout Chicago

In this section, we use our estimated structural model to conduct a neighborhood-level analysis of how rents and racial characteristics changed as a result of public housing demolitions. Our analysis extends on the descriptive patterns documented in Section 4.1. Previously, we documented that census tracts with more demolitions experienced larger changes in housing market prices and demographics between 2000 and 2010. However, these descriptive results do not disentangle all effects generated by demolitions. Perhaps most importantly, the previous analysis cannot quantify the spatial equilibrium effects that may generate spillovers in areas not directly affected by demolitions.

Theoretically, our model suggests that there will be significant heterogeneity in the effects of public housing demolitions across neighborhoods. Given the household preference estimates, we expect rents to go up in neighborhoods which had demolitions. For neighborhoods without demolitions, there are two potentially offsetting effects that arise through substitution across neighborhoods and equilibrium forces. First, demand might shift toward neighborhoods that directly

---

[36]We calculate these homeownership rates using data on households in the Chicago metropolitan area using 2008–2012 ACS data.

experienced demolitions, so the relative value of neighborhoods without demolitions could fall, which we refer to as the cross-demolition elasticity effect. Second, the increase in housing prices in neighborhoods with demolitions could increase demand for substitutes of those neighborhoods, which we refer to as the cross-price elasticity effect. Which effect dominates depends on the estimated willingness to pay of households as well as on the housing supply elasticity.

To estimate the direct and spillover effects of public housing demolitions, we compare tract-level variables in the actual scenario for 2010, where demolitions occurred, to the estimated counterfactual scenario in which there are no demolitions. We begin our analysis by focusing on tract-level impacts on rental prices. Figure 9 provides separate histograms for tracts with and without public housing demolitions. Tracts with a demolition saw an average rent increase of 13.8 percent (Panel A). The distribution of changes for these tracts exhibits a fat right tail. These areas with the largest rent changes had the most extensive number of demolitions. The 13.8 percent effect for tracts with demolitions is over 7 times larger than the average rent increase of 1.9 percent in tracts without demolitions (Panel B). The fact that prices increase in neighborhoods without demolitions implies that the cross-price effect dominates the cross-demolition effect in our empirical analysis. Moreover, there is substantial heterogeneity in the size of the rent increase, reflecting variation in the extent of demolitions in a neighborhood and the desirability of the neighborhood on other dimensions.

To more clearly demonstrate how rent changes vary with distance to public housing demolitions, Panel C of Figure 9 displays a map of the tract-level changes in log rents. The darkest shaded areas on the map again indicate that neighborhoods with public housing demolitions saw the largest increases in rents. As expected given the disadvantaged nature of public housing neighborhoods, these areas experiencing large increases are those that would have had the lowest rent in the 2010 no-demolition counterfactual.[37] Also apparent is that neighborhoods that are relatively close to demolition areas—other tracts in the south and west sides of the city—experienced moderate increases in rents. Appendix Figure A.3 quantifies this by illustrating the relationship between the tract-level price effect and distance to public housing demolitions. These results show that the average rent increase for neighborhoods without demolitions is about 2.5 percent for tracts that are within 0.1 miles of a demolition site, but only 1 percent for neighborhoods that are 25 miles away (near the border of Cook County).

In addition to studying rents, we also examine how demolitions impacted neighborhood demographic composition. Figure 10 plots the demolition-induced change in the share of households that are not poor and White against the change in the share of households that are poor and Black

---

[37]In Appendix Figure A.2, we plot the tract-level change in log rents (y-axis) against the estimated rent in each tract from the 2010 no-demolition counterfactual (x-axis). These results clearly show that demolitions had the largest impact on the lowest price neighborhoods.

for each tract in Chicago. The slope coefficient is precisely estimated and implies that areas where demolitions caused a 1 percent decrease in the poor Black household share experienced a 0.53 percent increase in the non-poor White household share on average. This result is consistent with the descriptive evidence in Figure 3 and further indicates that public housing demolitions were followed by neighborhood change. While there is little consensus on whether gentrification and more general forms of neighborhood change have led to decreases in the welfare of poor minority households (e.g. Vigdor, 2002; Brummet and Reed, 2021), our structural model implies that public housing demolitions did lead to welfare declines for these groups.

Our results underscore the benefits of using a structural model to study the consequences of demolitions. In particular, we build on prior studies that use reduced form approaches and find public housing demolitions increased property values by 9 to 20 percent in directly targeted areas (Brown, 2009; Zielenbach and Voith, 2010; Blanco and Neri, 2021). While our estimated direct impacts are in line with prior findings, our model-based approach allows us to estimate how demolitions affect equilibrium outcomes in *each* neighborhood in Chicago. This allows us to provide new evidence that equilibrium spillovers driven by choice substitution have positive impacts on tracts throughout Chicago. As summarized in Table 3, our estimates imply that 74 percent of the aggregate increase in rents comes from neighborhoods without public housing demolitions.[38]

9   Welfare Impacts Under Alternative Housing Policies

A natural consideration is how additional housing policy responses influence the welfare consequences of public housing demolitions. In this section, we use our structural model to study the effects of two types of interventions that might mitigate the disparate impacts of demolitions. First, we examine the importance of relaxing restrictions on building and improvements in the regulatory environment. In the context of our framework, we approximate this type of policy response by studying how welfare depends on the housing supply elasticity (c.f. Gyourko et al., 2008; Saiz, 2010). Second, we study how the scale of redevelopment in public housing areas matters by estimating counterfactuals in which we assume additional market rate units are created in neighborhoods which featured demolitions.

We begin our analysis by calculating the welfare impacts of demolitions under scenarios where we vary the housing supply elasticity, ψ, in equation (4). Our analysis considers supply elasticities that range between 0 and 0.70. This upper bound is based on the maximum estimate in Baum-Snow and Han (2021). Theoretically, more elastic housing supply would result in additional housing units in neighborhoods that become more attractive after demolitions. As a result, a higher housing

---

[38]Put differently, the 5 percent of neighborhoods with a public housing demolition account for 26 percent of the city-wide housing price increase.

supply elasticity would reduce the positive price impacts that particularly reduce the welfare of poor households.

Panel A of Figure 11 displays the average impact on log rents due to demolitions for different assumed values of the housing supply elasticity by groupings of neighborhoods. We focus on rents given that our decomposition analysis demonstrates their central role in driving welfare impacts. Consistent with our neighborhood-level analysis in Section 8, we find that demolitions have positive average effects on rents for all types of neighborhoods that we consider. As the housing supply elasticity increases, the red line (circle marker) shows that there are particularly large declines in the effects of demolitions on neighborhoods directly receiving demolitions. For example, the direct effect of demolitions is 13.8 percent in our baseline specification when the housing supply elasticity is assumed to be 0.163. Increasing the housing supply elasticity to 0.45—approximately equal to the average elasticity for Youngstown, Ohio or Gary, Indiana estimated in Baum-Snow and Han (2021)—reduces the effect on rents by more than half to 6.6 percent. The blue line (square marker) shows that the qualitative patterns in the neighborhoods indirectly impacted are similar but muted relative to the neighborhoods where demolitions occur.

Panel B of Figure 11 shows that increases in the housing supply elasticity and the associated reduction in demolition-induced rent increases have heterogeneous effects on welfare across racial and income groups. Poor Black and Hispanic households (blue, solid square and green, solid triangle markers, respectively) benefit the most from scenarios that have larger housing supply responses. For example, increasing the housing supply elasticity from our baseline value of 0.163 to 0.45 reduces the negative impact of demolitions from -$75 to $10 in the rent equivalent for poor Black households. While greater housing supply responses improve outcomes for minority households, both poor and non-poor White households have reduced welfare gains. This stems from the fact that homeowners—who constitute the majority of the White population even among poor households—have lower rental income when the elasticity of housing supply is larger.[39] Overall, the pattern of results suggests that a higher housing supply elasticity lessens racial disparities in the effects of demolitions although welfare losses are only eliminated when the assumed elasticity is at the upper range of the estimates from Baum-Snow and Han (2021).

Next, we consider a more-targeted alternative to pursuing policies that seek to broadly increase housing supply responses: generating greater redevelopment in neighborhoods where public housing was demolished. Specifically, we consider a scenario where the city government targets areas with demolitions and creates additional market-rate housing units. This exercise differs from our main analysis which incorporates the observed amount of redevelopment in 2010. Notably, our exercise of simulating *additional* housing redevelopment is also motivated by the fact that such a policy is feasible given the high vacancy rates that persisted in former public housing neighbor-

---

[39]Appendix Figure A.4 provides welfare estimates separately for renters and homeowners.

hoods (see Appendix Table A.1).[40] Allowing for greater redevelopment in the form of publicly-built market-rate housing units serves to increase the supply of housing in neighborhoods where poor minority households tended to live before demolition.

To understand the impact of expanding redevelopment, Figure 12 shows how the demolition effects on rents (Panel A) and welfare (Panel B) vary with the scale of redevelopment. We characterize redevelopment in terms of the share of total public housing that was destroyed. At the maximum of 0.5, we assume that the government constructs additional market-rate housing in demolition neighborhoods that replaces 50% of the units that we observe as having been destroyed in 2010. This upper bound is motivated by the fact that about 50% of former public housing sites were not re-developed for residential or commercial uses (Appendix Table A.1).

In Panel A, we find that expanding redevelopment in public housing areas has large impacts on reducing the effects of demolitions on rents. Notably, the results show that replacing approximately 30 percent of the destroyed public housing stock with market rate housing eliminates the effects of public housing demolition on rents in neighborhoods that experienced demolitions (red line, circle marker). At this level of expanded redevelopment, positive impacts on rents are still present in neighborhoods that did not have demolitions (average increase: 1.4 percent), although this change is about 25 percent lower than the increase without additional redevelopment.

The results in Panel B show that the impact of demolition on welfare improves for most groups with the scale of redevelopment. For all types of households, redeveloping 20 percent of the destroyed housing is a sufficient intervention that results in welfare gains associated with demolition. At even higher levels of redevelopment, the rent equivalent welfare impacts of demolition for minorities begin to equal or exceed the positive impacts for poor and non-poor White households, who are made worse off by the decrease in the value of their homes (and thereby the reductions in their rental income).

A comparison of the results in Figures 11 and 12 highlights several key differences from the alternative housing policies. Increasing the scale of redevelopment does relatively more to reduce the effects of demolitions on rental prices in targeted neighborhoods. In line with this result, public housing demolitions in scenarios which feature high levels of redevelopment have positive impacts on rent equivalent welfare for all groups.

Why does redevelopment reverse the negative impacts of public housing demolition for poor and minority households? A key difference is that a policy of redevelopment increases the supply of housing in neighborhoods that are ex-ante cheaper—neighborhoods which featured public housing demolition have an average monthly rental price of $644 in 2000 which stands at the 6th

---

[40]In this scenario, we allow for this government redevelopment to crowd out private housing construction. The response of the latter is always summarized by the supply function in equation (4). Note that we do not require the local government to pay for redevelopment, so this exercise is best viewed as representing the consequences of expanding the federally-funded HOPE VI program to include more extensive redevelopment efforts.

percentile of the city-wide distribution. As a result, redevelopment leads to larger housing price declines at the low end of the price distribution. Given that minorities and poor households are more sensitive to housing prices, they tend to live in cheaper areas. Overall, redevelopment in neighborhoods with public housing demolitions has larger distributional implications by indirectly targeting minorities through their location choices.

10 Conclusion

This paper provides new evidence on the welfare consequences of urban renewal programs by studying federally-funded public housing demolitions in Chicago. As noted in prior literature, these demolitions led to lasting changes in the housing market and demographic composition of targeted neighborhoods. We use a structural approach to quantify how these changes shaped welfare and study distributional considerations across racial and income groups.

Our main finding is that demolitions had disparate impacts and generated large welfare improvements for White households alongside welfare losses for low-income minority households. The unequal effects of demolitions arise from several forces. First, while all households benefit from the destruction of public housing, reductions in the racial minority share in targeted neighborhoods and subsequent resorting generate large gains for White households and losses for Black and Hispanic households. Second, increases in rental prices further exacerbate racial inequality in the effects of demolition because White households benefited from this price appreciation due to their high rates of home ownership.

Overall, the disparate impacts on welfare in our results highlight fundamental limitations of policies that aim to revitalize neighborhoods and benefit lower-income households. While these types of urban dynamics have been discussed qualitatively in prior work (Glaeser and Gottlieb, 2008; Neumark and Simpson, 2015), our structural approach allows us to break new ground by explicitly quantifying these effects in the context of one of the largest place-based programs pursued in the U.S. Our findings should be relevant more broadly to the extent that other policies generate large-scale resorting and preferences over racial composition and price sensitivity in other settings are similar to those in Chicago.

Finally, a key policy implication of our results is that redevelopment can potentially play a key role in shaping welfare impacts of urban renewal programs such as public housing demolition. We find that moderate increases in the scale of housing redevelopment in areas targeted by demolition reverse the negative impacts of public housing demolition and allow all racial and income groups to benefit. This finding shapes historical perspectives of U.S. housing policies during the past three decades. The welfare impacts of public housing demolitions in Chicago may have been more positive if authorities had engaged in more intensive redevelopment efforts. More broadly, major

U.S. cities such as Atlanta and Washington, D.C. also received substantial HOPE VI funding. The well-documented lack of redevelopment in many of these cities (Vale et al., 2018) may have muted the welfare benefits of public housing demolition for minority and lower-income residents.

References

Aliprantis, Dionissi and Hartley, Daniel. (2015). 'Blowing It Up and Knocking It Down: The Local and City-wide Effects of Demolishing High Concentration Public Housing on Crime', Journal of Urban Economics 88, 67–81.

Almagro, Milena and Domınguez-Iino, Tomás. (2019). 'Location Sorting and Endogenous Amenities: Evidence from Amsterdam'.

Balboni, Clare, Bryan, Gharad, Morten, Melanie and Siddiqi, Bilal. (2020). 'Transportation, Gentrification, and Urban mobility: The Inequality effects of place-based policies', Preliminary Draft p. 3.

Baum-Snow, Nathaniel and Han, Lu. (2021), The Microgeography of Housing Supply.

Baum-Snow, Nathaniel and Marion, Justin. (2009). 'The Effects of Low Income Housing Tax Credit Developments on Neighborhoods', Journal of Public Economics 93(5–6), 654–666.

Bayer, Patrick, Casey, Marcus, McCartney, W Ben, Orellana-Li, John and Zhang, Calvin. (2022). 'Distinguishing Causes of Neighborhood Racial Change: A Nearest Neighbor Design'.

Bayer, Patrick, Ferreira, Fernando and McMillan, Robert. (2007). 'A Unified Framework for Measuring Preferences for Schools and Neighborhoods', Journal of Political Economy 115(4), 588–638.

Bayer, Patrick, McMillan, Robert, Murphy, Alvin and Timmins, Christopher. (2016). 'A Dynamic Model of Demand for Houses and Neighborhoods', Econometrica 84(3), 893–942.

Bayer, Patrick and Timmins, Christopher. (2005). 'On the Equilibrium Properties of Locational Sorting Models', Journal of Urban Economics 57(3), 462–477.

Berry, Steven, Levinsohn, James and Pakes, Ariel. (1995). 'Automobile Prices in Market Equilibrium', Econometrica: Journal of the Econometric Society pp. 841–890.

Berry, Steven T. (1994). 'Estimating Discrete-choice Models of Product Differentiation', The RAND Journal of Economics pp. 242–262.

Bittle, Jake, Kapur, Srishti and Mithani, Jasmine. (2017). 'Redeveloping the State Street Corridor', South Side Weekly p. 15.

Blanco, Hector. (2021), Pecuniary Effects of Public Housing Demolitions: Evidence from Chicago, Technical report, Working Paper.

Blanco, Hector and Neri, Lorenzo. (2021). 'Knocking it Down and Mixing it Up: The Impact of Public Housing Regenerations'.

Brown, LaTanya N. (2009). 'HOPE VI: An Analysis to Determine the HOPE VI Program's Influence on Home Sales', Community Development 40(1), 54–63.

Brummet, Quentin and Reed, Davin. (2021), 'The Effects of Gentrification on Incumbent Residents'.

Busso, Matias, Gregory, Jesse and Kline, Patrick. (2013). 'Assessing the Incidence and Efficiency of a Prominent Place Based Policy', American Economic Review 103(2), 897–947.

Caetano, Gregorio and Maheshri, Vikram. (2021), A Unified Empirical Framework to Study Segregation.

Christensen, Peter and Timmins, Christopher. (2021). 'Sorting or Steering: The Effects of Housing Discrimination on Neighborhood Choice', Working Paper .

Chyn, Eric. (2018). 'Moved to Opportunity: The Long-Run Effects of Public Housing Demolition on Children', The American Economic Review 108(10), 3028–3056.

Collins, William J. and Shester, Katharine L. (2013). 'Slum Clearance and Urban Renewal in the United States', American Economic Journal: Applied Economics 5(1), 239–273. Publisher: American Economic Association.

Couture, Victor, Gaubert, Cecile, Handbury, Jessie and Hurst, Erik. (2019), Income Growth and the Distributional Effects of Urban Spatial Sorting, Technical report, National Bureau of Economic Research.

Couture, Victor and Handbury, Jessie. (2020). 'Urban Revival in America', Journal of Urban Economics 119, 103267.

Davis, Morris A, Gregory, Jesse and Hartley, Daniel A. (2019), The Long-Run Effects of Low-Income Housing on Neighborhood Composition.

Davis, Morris A, Gregory, Jesse, Hartley, Daniel A and Tan, Kegon TK. (2021). 'Neighborhood Effects and Housing Vouchers', Quantitative Economics 12(4), 1307–1346.

Diamond, Rebecca. (2016). 'The Determinants and Welfare Implications of US Workers' Diverging Location Choices by Skill: 1980-2000', American Economic Review 106(3), 479–524.

Diamond, Rebecca and McQuade, Tim. (2019). 'Who Wants Affordable Housing in Their Backyard? An Equilibrium Analysis of Low-Income Property Development', Journal of Political Economy 127(3), 1063–1117.

Diamond, Rebecca, McQuade, Timothy and Qian, Franklin. (2018), The Effects of Rent Control Expansion on Tenants, Landlords, and Inequality: Evidence from San Francisco, Working Paper 24181, National Bureau of Economic Research. Series: Working Paper Series.

Diamond, Rebecca, McQuade, Tim and Qian, Franklin. (2019). 'The Effects of Rent Control Expansion on Tenants, Landlords, and Inequality: Evidence from San Francisco', American Economic Review 109(9), 3365–3394.

Dumke, Mick. (2017). 'Years late, with big gaps, CHA nears end of housing 'transformation'', Chicago Sun Times p. 11.

Fu, Chao and Gregory, Jesse. (2019). 'Estimation of an Equilibrium Model With Externalities: Post-Disaster Neighborhood Rebuilding', Econometrica 87(2), 387–421.

Galiani, Sebastian, Murphy, Alvin and Pantano, Juan. (2015). 'Estimating Neighborhood Choice Models: Lessons from a Housing Assistance Experiment', American Economic Review 105(11), 3385–3415.

Gechter, Michael and Tsivanidis, N. (2020). 'Spatial Spillovers From Urban Renewal: Evidence from the Mumbai Mills Redevelopment', Penn State University (mimeo) .

Glaeser, Edward L. and Gottlieb, Joshua D. (2008). 'The Economics of Place-Making Policies', Brookings Papers on Economic Activity 2008, 155–239. Publisher: Brookings Institution Press. **URL:** *http://www.jstor.org/stable/27561617*

Guerrieri, Veronica, Hartley, Daniel and Hurst, Erik. (2013). 'Endogenous Gentrification and Housing Price Dynamics', Journal of Public Economics 100, 45–60.

Gyourko, Joseph, Saiz, Albert and Summers, Anita. (2008). 'A New Measure of the Local Regulatory Environment for Housing Markets: The Wharton Residential Land Use Regulatory Index', Urban Studies 45(3), 693–729.

Harari, Mariaflavia and Wong, Maisy. (2018), Slum Upgrading and Long-Run Urban Development: Evidence from Indonesia, in '2018 Meeting Papers. Society for Economic Dynamics'.

Hunt, D. Bradford. (2009), Blueprint for Disaster: The Unraveling of Chicago Public Housing, University of Chicago Press.

Jacob, Brian A. (2004). 'Public Housing, Housing Vouchers, and Student Achievement: Evidence from Public Housing Demolitions in Chicago', The American Economic Review 94(1), 233–258.

Khanna, Gaurav, Medina, Carlos, Nyshadham, Anant, Ramos-Menchelli, Daniel, Tamayo, Jorge and Tiew, Audrey. (2022). 'Spatial Mobility, Economic Opportunity, and Crime', Preliminary Draft .

Kline, Patrick and Moretti, Enrico. (2014). 'People, Places, and Public Policy: Some Simple Welfare Economics of Local Economic Development Programs', Annual Review of Economics 6(1), 629–662.

McFadden, Daniel. (1974). 'The Measurement of Urban Travel Demand', Journal of public economics 3(4), 303–328.

Neumark, David and Simpson, Helen. (2015), Place-Based Policies, in 'Handbook of Regional and Urban Economics', Vol. 5, Elsevier, pp. 1197–1287.

**Nevo, Aviv**. (2001). 'Measuring Market Power in the Ready-to-eat Cereal Industry', Econometrica 69(2), 307–342.

Popkin, Susan J, Gwiasda, Victoria, Olson, Lynn, Rosenbaum, Dennis and Buron, Larry. (2000), The Hidden War: Crime and the Tragedy of Public Housing in Chicago, Rutgers University Press, New Brunswick, NJ.

Qian, Franklin and Tan, Rose. (2021), The Effects of High-skilled Firm Entry on Incumbent Residents, Stanford Institute for Economic Policy Research (SIEPR).

Rossi-Hansberg, Esteban, Sarte, Pierre-Daniel and Owens, Raymond. (2010). 'Housing Externalities', Journal of Political Economy 118(3), 485–535.

Saiz, Albert. (2010). 'The Geographic Determinants of Housing Supply', Quarterly Journal of Economics 125(3), 1253–1296.

Sandler, Danielle H. (2017). 'Externalities of Public Housing: The Effect of Public Housing Demolitions on Local Crime', Regional Science and Urban Economics 62, 24–35.

Scott, Paul T. (2013). 'Dynamic Discrete Choice Estimation of Agricultural Land Use', Toulouse School of Economics Working Paper 526.

Story, Louise. (2012). 'As Companies Seek Tax Deals, Governments Pay High Price', New York Times .

Su, Yichen. (2022). 'The Rising Value of Time and the Origin of Urban Gentrification', American Economic Journal: Economic Policy 14(1), 402–39.

Tach, Laura and Emory, Allison Dwyer. (2017). 'Public Housing Redevelopment, Neighborhood Change, and the Restructuring of Urban Inequality', American Journal of Sociology 123(3), 686–739.

U.S. National Commission on Severely Distressed Public Housing. (1992), Final Report of the National Commission on Severely Distressed Public Housing, Technical report, Washington, D.C.

Vale, Lawrence J., Shamsuddin, Shomon and Kelly, Nicholas. (2018). 'Broken Promises or Selective Memory Planning? A National Picture of HOPE VI Plans and Realities', Housing Policy Debate pp. 1–24.

Vigdor, Jacob L. (2002). 'Does Gentrification Harm the Poor?', Brookings-Wharton Papers on Urban Affairs pp. 133–182.

Wong, Maisy. (2013). 'Estimating Ethnic Preferences Using Ethnic Housing Quotas in Singapore', Review of Economic Studies 80(3), 1178–1214.

Zielenbach, Sean and Voith, Richard. (2010). 'HOPE VI and Neighborhood Economic development: The Importance of Local Market Dynamics', Cityscape pp. 99–131.

Figure 1: Time Series of Public Housing Demolitions in Chicago

(a) Chicago
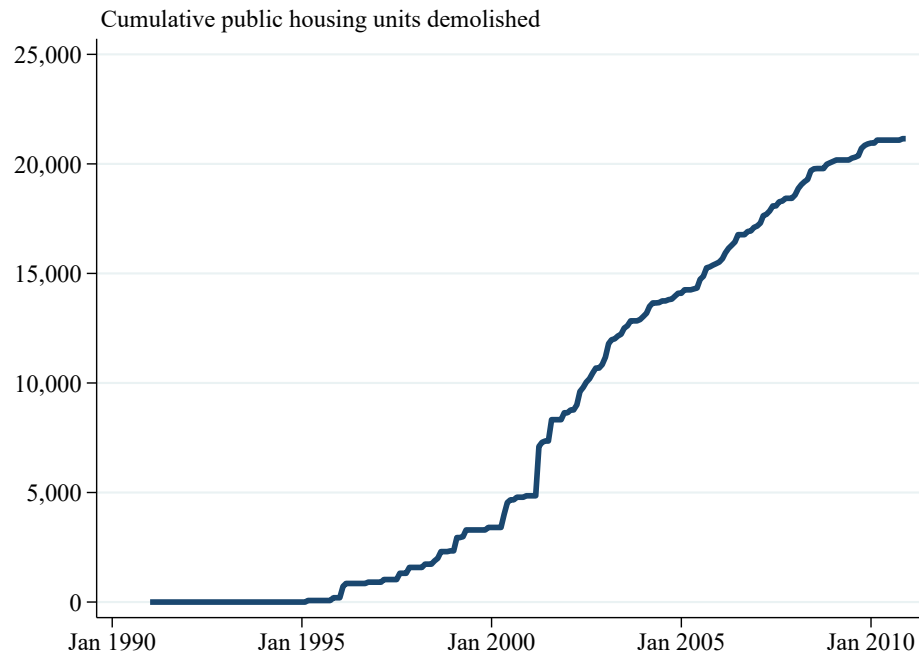
(b) By Tract
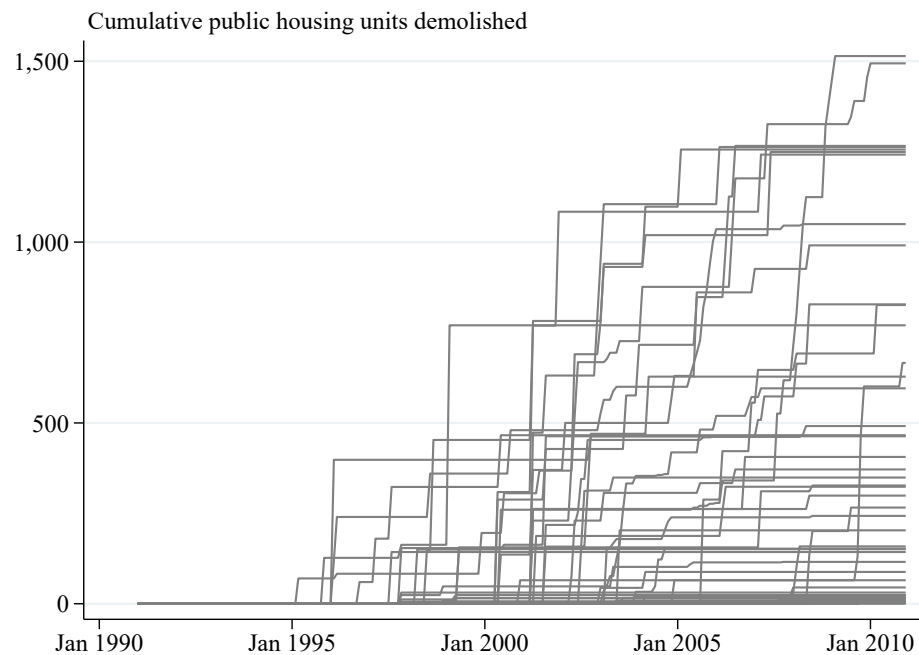
Notes: Panel A displays the cumulative number of public housing units that were demolished in Chicago between 1995 and 2010. Panel B displays results separately for each of the 59 census tracts that experienced a demolition.

Source: Authors' calculations using data from the Chicago Housing Authority.

Figure 2: Spatial Variation in Public Housing Demolitions in Chicago

(a) Cumulative Demolitions by Tract

Legend:
- [1,3]
- (3,8]
- (8,21]
- (21,55]
- (55,152]
- (152,283]
- (283,406]
- (406,647]
- (647,1050]
- (1050,1514]
- No demolitions
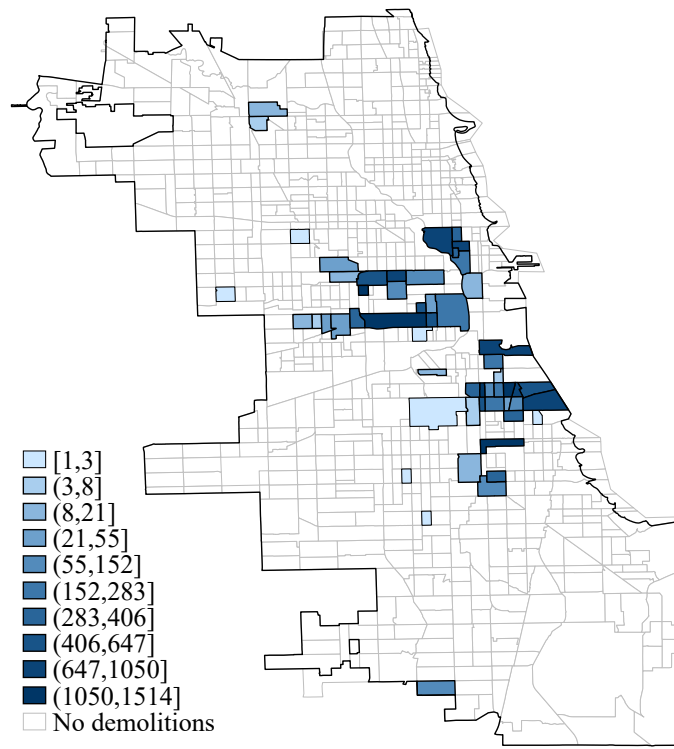
(b) Distribution of Cumulative Demolitions as Share of 1990 Housing Stock
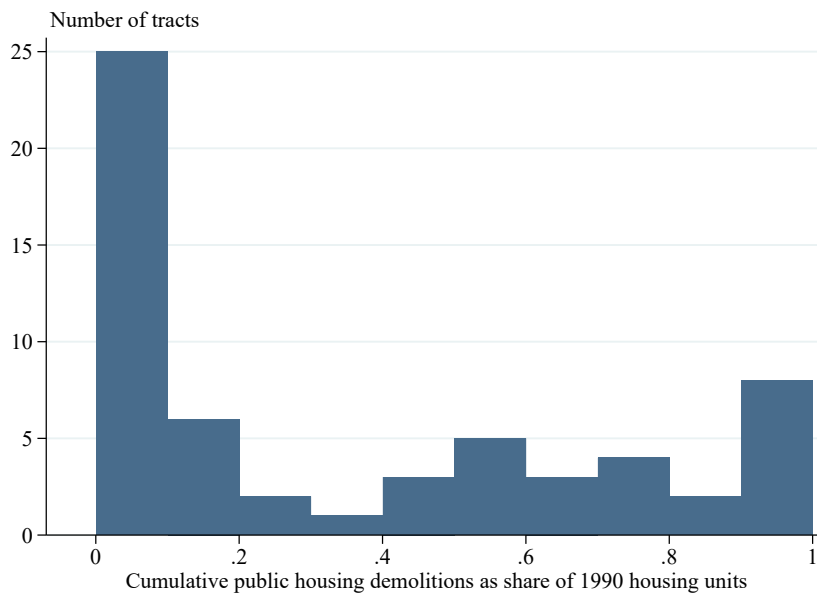
Notes: Panel A displays the cumulative number of public housing units that were demolished in each census tract between 1995 and 2010. Panel B displays the cumulative number of demolitions as a share of the number of occupied housing units in 1990 for tracts that experienced a demolition. We winsorize this variable from above at 1 for 7 tracts, but results are not sensitive to this choice. The width of each bar in Panel B is 0.1.

Source: Authors' calculations using data from the Chicago Housing Authority.

Figure 3: Changes in Demographics and Public Housing Demolitions, 2000–2010 and 2000–2016

(a) Non-Hispanic White Population Share

(b) Black Population Share

(c) Hispanic Population Share

(d) Log Median Household Income
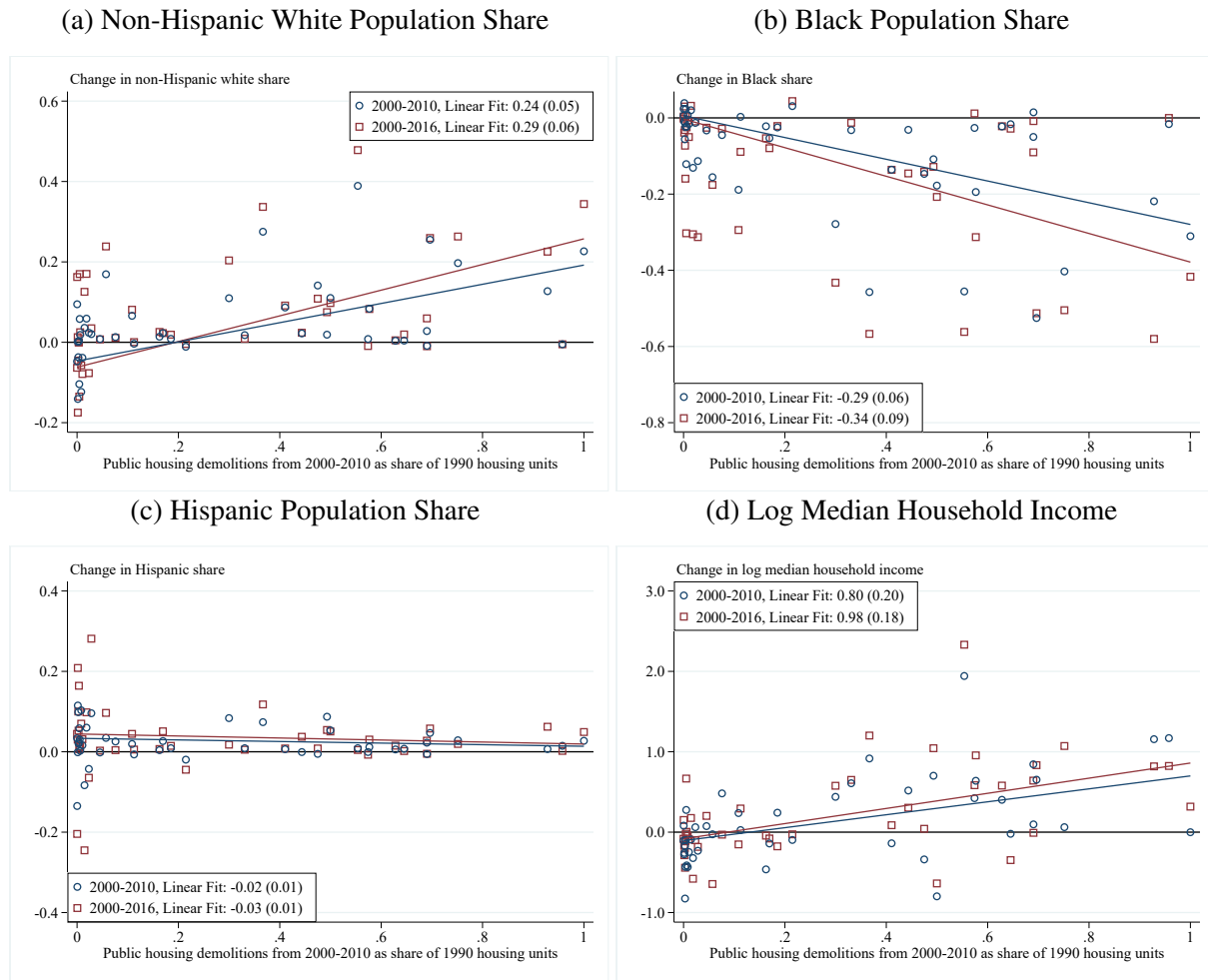
Notes: This figure plots the change in neighborhood characteristics against the cumulative number of public housing units demolished from 2000–2010 as a share of the number of occupied housing units in 1990. We winsorize the public housing demolition share variable from above at 1 for 3 tracts. Each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 4: Changes in Housing Market Characteristics and Public Housing Demolitions, 2000–2010 and 2000–2016

(a) Log Median Rent

(b) Log Median House Value

(c) Share of Housing Units 0–10 Years Old

(d) Share of Housing Units 31+ Years Old
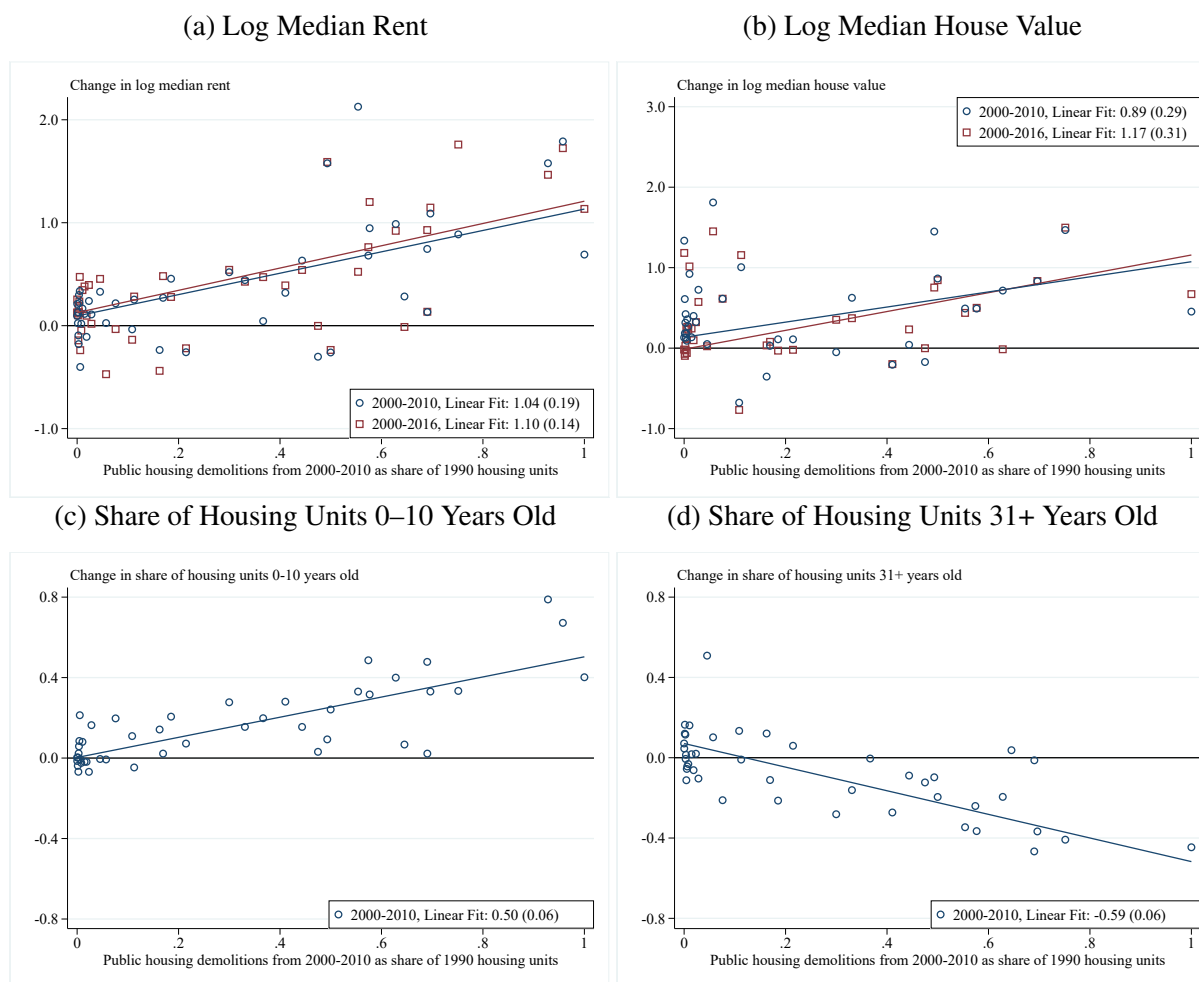
Notes: See notes to Figure 3. Panels C and D do not contain 2000–2016 changes because the available ACS data for 2016 do not allow us to construct the same housing unit age bins.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 5: 1990–2000 Changes in Rents and Housing Values Compared to 2000–2010 Changes in Public Housing Demolitions and Instrumental Variable Predicted Change in Rents

(a) Change in Log Rent Against Demolitions

(b) Change in Log Rent Against Instrument

(c) Change in Log House Value Against Demolitions

(d) Change in Log House Value Against Instrument
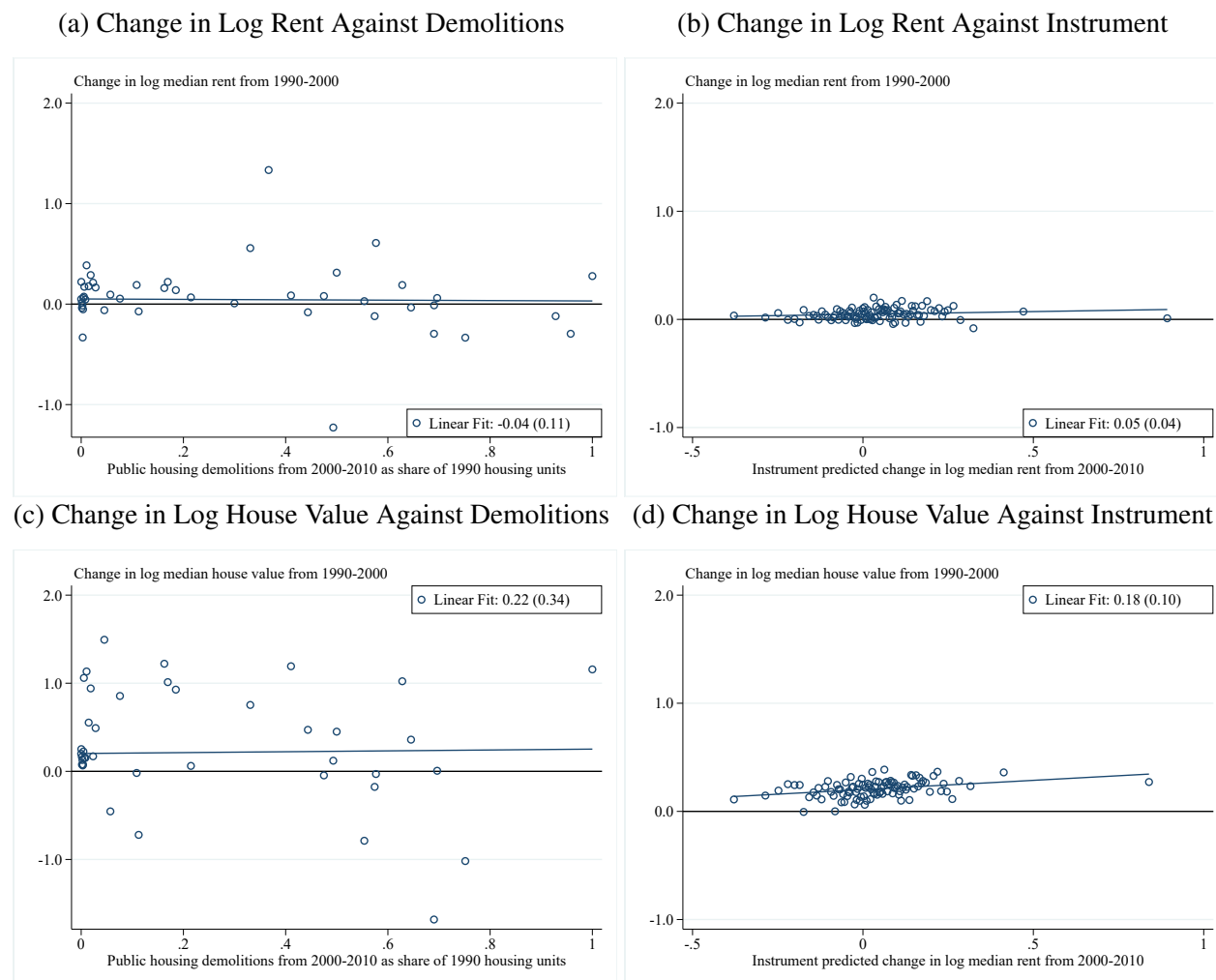
Notes: This figure plots the change in log median rent and log median house value from 1990–2000 against the number of public housing units demolished from 2000–2010 as a share of the number of occupied housing units in 1990 (Panels A and C) and the predicted change in log rent from 2000–2010 based on our instrumental variable procedure (Panels B and D). In Panels A and C, each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition. In Panels B and D, each dot represents the average change for each percentile of the instrument-predicted change in log rent.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 6: Assessing In-Sample Fit of Structural Model Using Rent Data

(a) Using Implied Supply, 2000

Slope: 1.017 (0.003), R2: 0.99

(b) Using Implied Supply, 2010

Slope: 1.004 (0.003), R2: 0.99

(c) Using Observed Supply, 2000

Slope: 1.044 (0.008), R2: 0.97

(d) Using Observed Supply, 2010

Slope: 1.039 (0.008), R2: 0.96
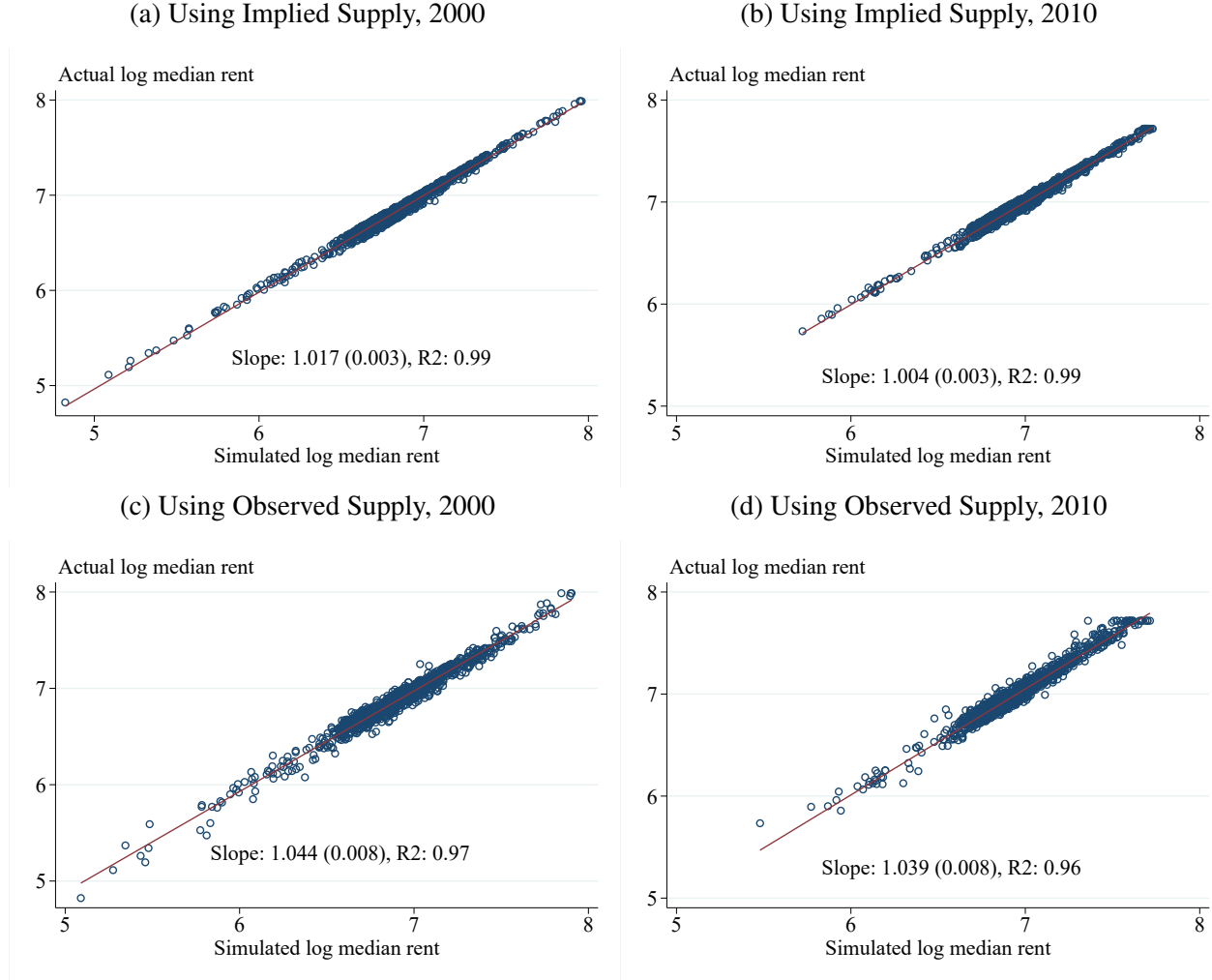
Notes: This figure plots actual log rents in census tracts against log rents that are implied by the model estimates where unobservable components of neighborhood quality are set to zero (i.e., $\tilde{\xi}_{jt}^k = 0$ for all k, j, and t). In Panels A and B, the number of housing units supplied is set to equal the number of housing units implied by the demand system. In Panels C and D, the number of housing units supplied is set to equal the observed number of housing units in census/ACS data (smoothed across tracts, to be consistent with the smoothed choice probabilities).

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 7: Assessing Out-of-Sample Fit of Structural Model Using Rent Data

Slope: 0.957 (0.036), R2: 0.70
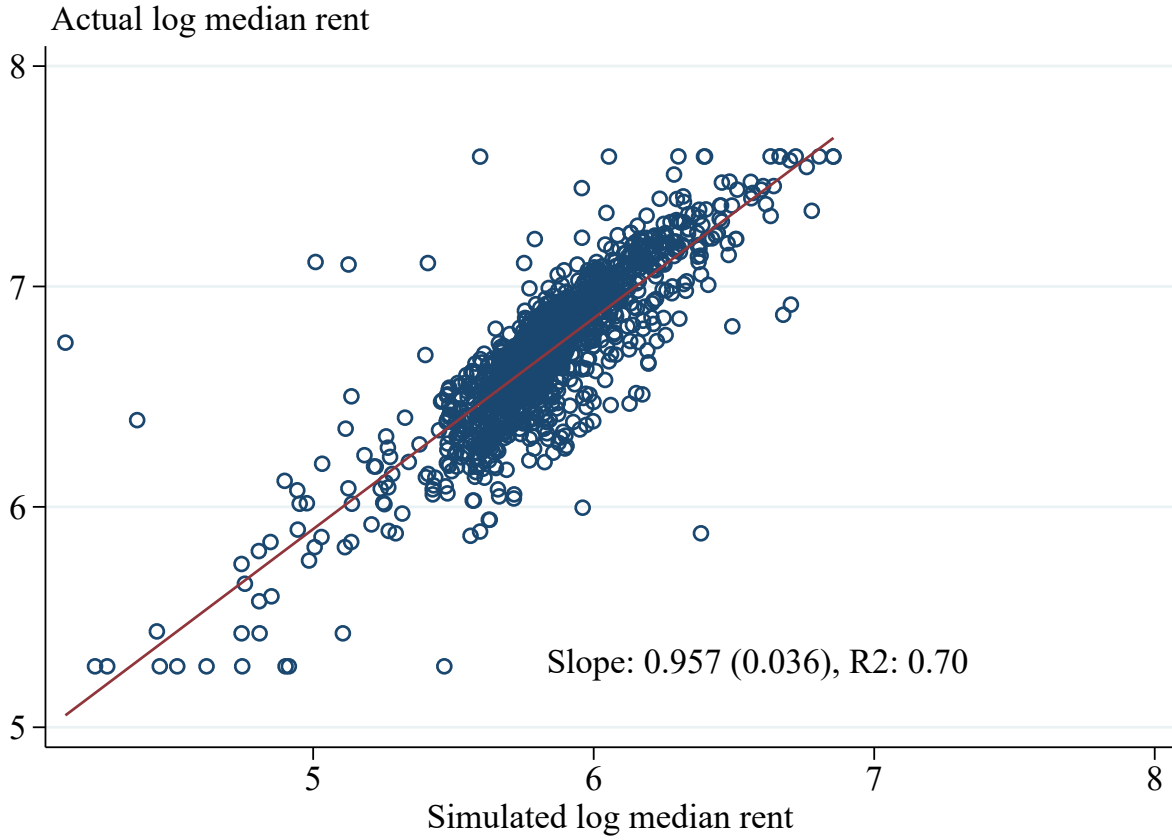
Notes: This figure plots actual log rents in 1990 in census tracts against log rents that are simulated by an out-of-sample procedure. In particular, we construct simulated rents for 1990 using the coefficients and tract fixed effects estimated using 2000–2010 data, exogenous observed neighborhood characteristics in 1990, and the assumption that the housing supply shifter, θ_{jt}, is the same in 1990 and 2000. We then solve for the endogenous variables using the equilibrium definition in equations (5)–(7).

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 8: Summary of Welfare Consequences of Public Housing Demolitions
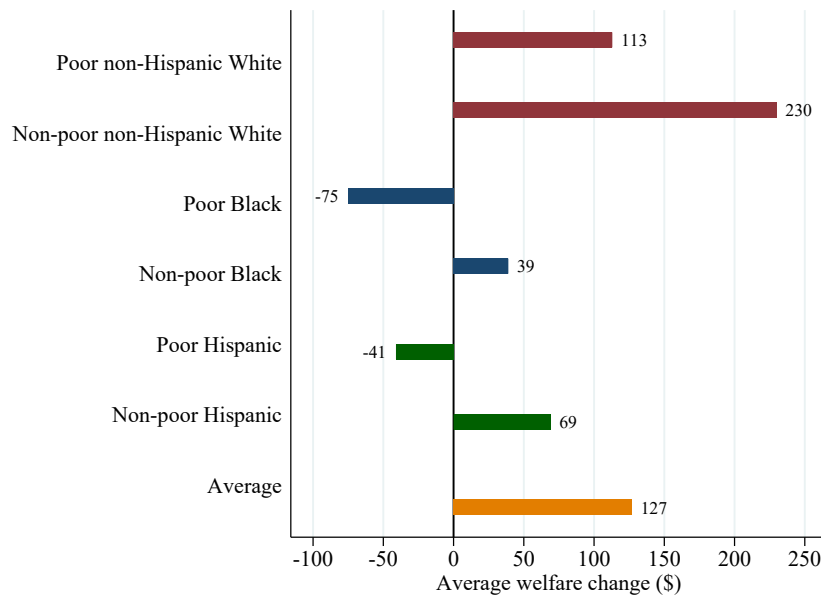
Notes: This figure reports the average change in each group's welfare due to public housing demolitions. We compare welfare under the actual state of the world in 2010 (where demolitions occurred) to a counterfactual version of 2010 in which the public housing share in each tract is held constant at its level in 2000. Welfare is expressed as the change in rents that would make households indifferent between the counterfactual and actual states of the world. This "rent equivalent" is normalized so that a positive value implies that demolitions lead to higher welfare. We construct the average rent equivalent as the population-weighted average of the group-specific rent equivalents for non-Hispanic White, Black, and Hispanic households.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Figure 9: Distribution of Tract-Level Rent Changes Due to Public Housing Demolitions

(a) Tracts with Public Housing Demolitions

Number of tracts: 57
Mean: 0.138, SD: 0.129
Min: 0.015, Max: 0.455

(b) Tracts without Public Housing Demolitions

Number of tracts: 1,183
Mean: 0.019, SD: 0.005
Min: -0.000, Max: 0.035

(c) All Tracts

[0.00,0.02]
(0.02,0.05]
(0.05,0.10]
(0.10,0.25]
(0.25,0.50]
No estimate
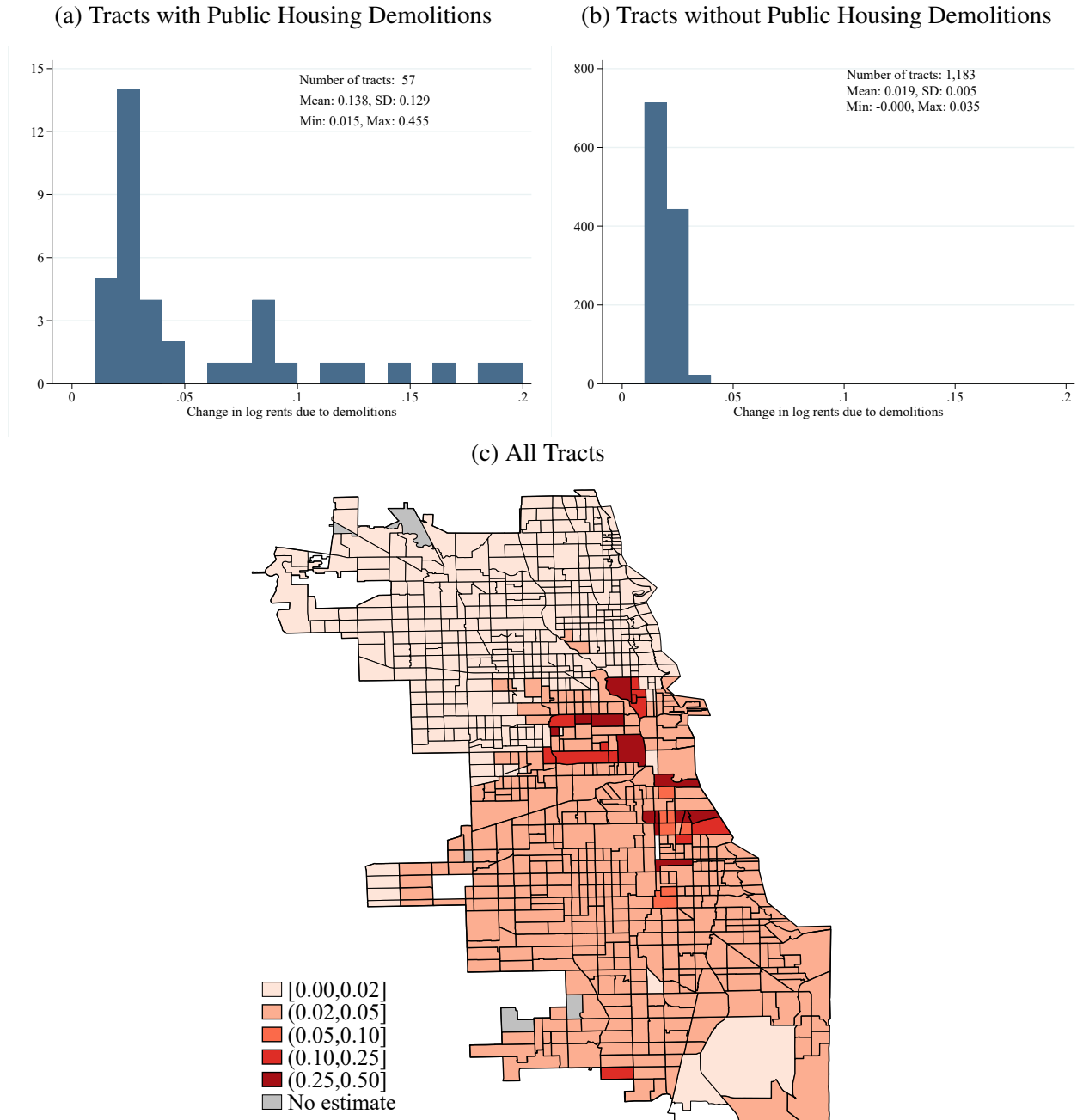
Notes: This figure displays the distribution of the change in log median rents due to public housing demolitions. We construct this change using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario where there are no demolitions. Panel A presents results for tracts where a public housing demolition occurred. Panel B presents results for tracts where a public housing demolition did not occur. In Panel A we omit 19 tracts where the change exceeds the included range. In Panel B we omit 1 tract with a change of -0.0002. The bin width is 0.01 in both panels. We calculate summary statistics using the number of households living in each tract as implied by the model. Panel C displays the tract-level change in log rents.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Figure 10: Public Housing Demolitions Increase Non-Poor White Share in Neighborhoods Where Poor Black Share Falls
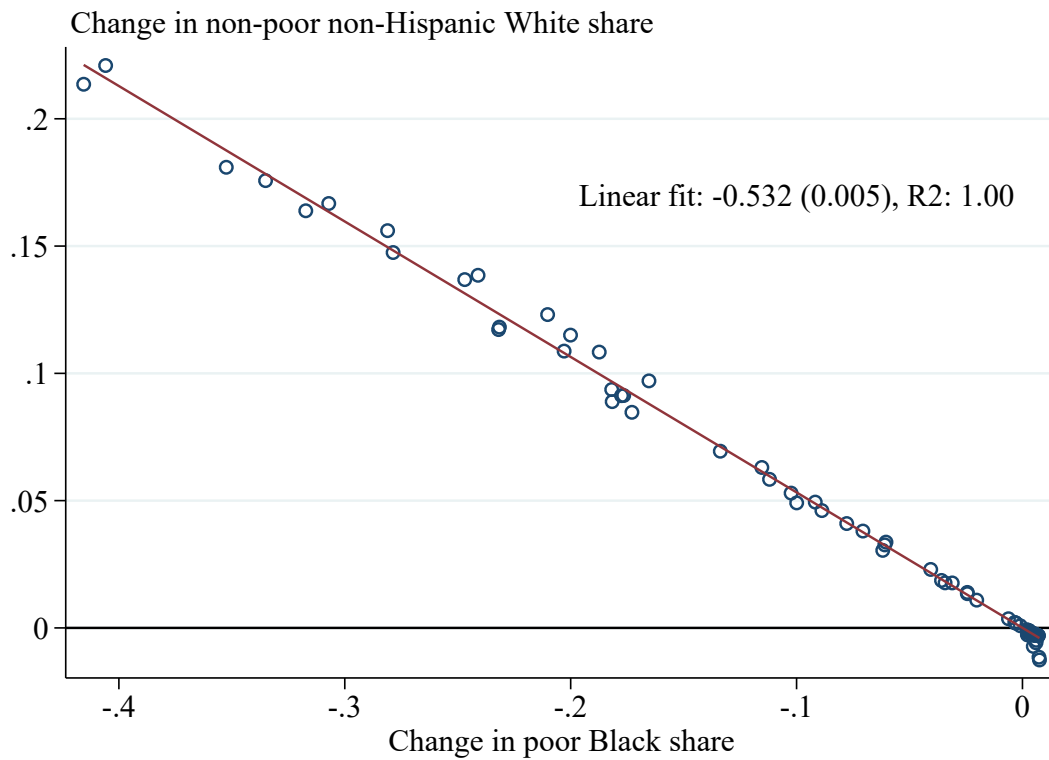
Notes: This figure displays the tract-level change in the share of households that are non-poor and White against the change in the share of households that are poor and Black. We construct these changes using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario where there are no demolitions.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Figure 11: Consequences of Public Housing Demolitions on Housing Prices and Welfare Under Alternative Housing Supply Elasticities

(a) Housing Price Consequences of Demolitions
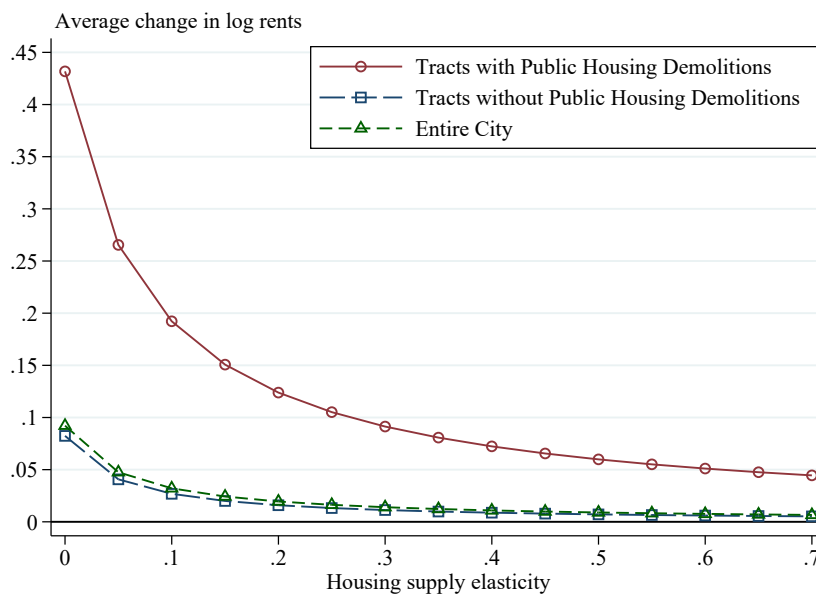
(b) Welfare Consequences of Demolitions
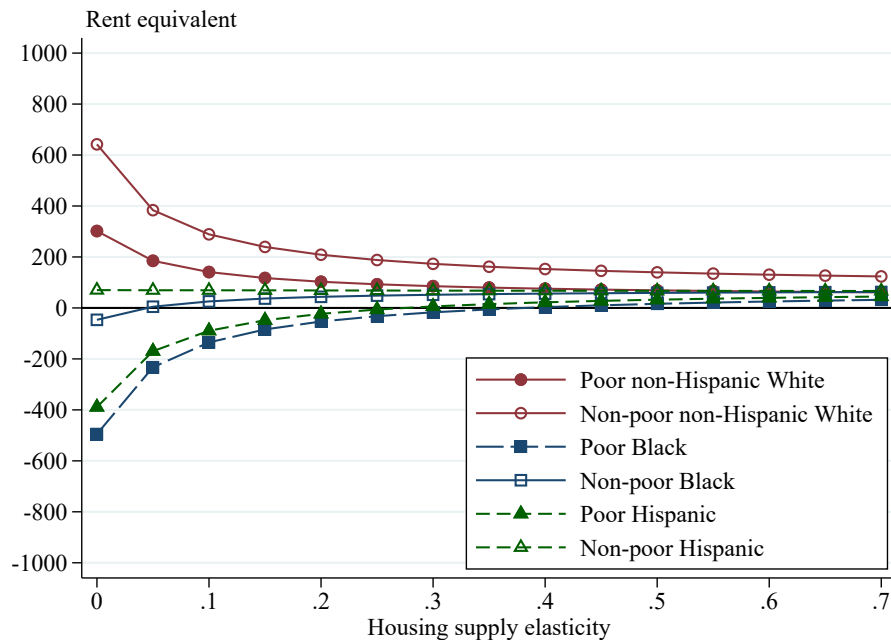
Notes: This figure displays outcomes from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Panel A shows the change in log rents due to demolitions under different assumptions about the housing supply elasticity. Panel B shows the rent equivalent welfare effect of public housing demolitions. Our baseline results are based on a housing supply elasticity of 0.163.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Figure 12: Consequences of Public Housing Demolitions on Housing Prices and Welfare Under Additional Redevelopment of Public Housing

(a) Housing Price Consequences of Demolitions
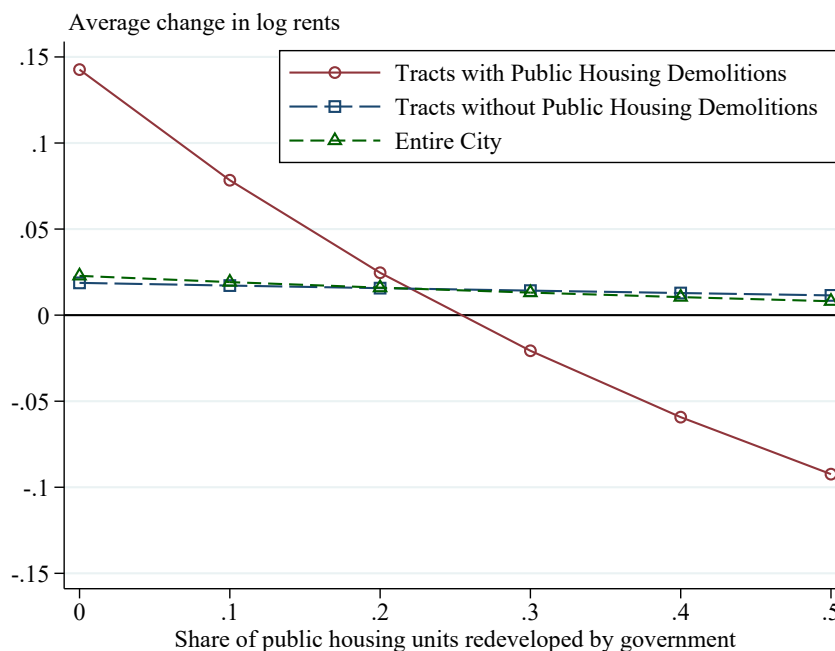
Average change in log rents

Share of public housing units redeveloped by government

- ○ — Tracts with Public Housing Demolitions
- ☐ — Tracts without Public Housing Demolitions
- △ — Entire City

(b) Welfare Consequences of Demolitions

Rent equivalent

Share of public housing units redeveloped by government

- ● — Poor non-Hispanic White
- ○ — Non-poor non-Hispanic White
- ■ — Poor Black
- ☐ — Non-poor Black
- ▲ — Poor Hispanic
- △ — Non-poor Hispanic

Notes: This figure displays outcomes from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. Panel A shows the change in log rents due to demolitions under different assumptions about the amount of additional redevelopment. Panel B shows the rent equivalent welfare effect of public housing demolitions. Our baseline results are given by 0 percent additional redevelopment.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Table 1: Instrumental Variable Estimates of Neighborhood Preference Parameters

| | Preference parameters for indicated group | | |
	Non-Hispanic White (1)	Black (2)	Hispanic (3)
Panel A: Poor Households			
Log median rent	-0.455***	-0.241***	-0.246***
	(0.0552)	(0.0333)	(0.0396)
Black population share	-0.165*	0.265***	0.273***
	(0.0929)	(0.0542)	(0.0573)
Hispanic population share	-0.132***	0.00195	0.307***
	(0.0365)	(0.0216)	(0.0283)
Log median household income	0.0886***	0.0353***	0.0128
	(0.0232)	(0.0127)	(0.0151)
Public housing units as a share of housing stock	-0.450***	-0.242***	-0.270***
	(0.107)	(0.0639)	(0.0733)
Panel B: Non-poor Households			
Log median rent	-0.0564***	-0.0368***	-0.103***
	(0.00908)	(0.0109)	(0.0263)
Black population share	-0.134***	0.220***	0.178***
	(0.0163)	(0.0212)	(0.0299)
Hispanic population share	-0.142***	0.0872***	0.261***
	(0.00712)	(0.00876)	(0.0199)
Log median household income	0.0217***	0.00978**	-0.00313
	(0.00387)	(0.00389)	(0.00916)
Public housing units as a share of housing stock	-0.0639***	-0.0848***	-0.127***
	(0.0153)	(0.0201)	(0.0415)
Specifications include:			
Year Fixed Effects	✓	✓	✓
Tract Fixed Effects	✓	✓	✓
Log median number of rooms	✓	✓	✓
Log median year built	✓	✓	✓
Homeownership share	✓	✓	✓
Land use variables	✓	✓	✓
Observations (tract-by-year)	2,480	2,480	2,480
Number of tracts	1,240	1,240	1,240

Notes: This table presents regression results of preference parameters for a static logit location choice model using household counts across census tracts in Cook County for 2000 and 2010. We estimate preference parameters separately by race/ethnicity and income group. Poor households have income below \$20,000, and non-poor households have income above \$20,000. Log median rent, Black and Hispanic population share, and log median income are instrumented following Bayer et al. (2007), where we take changes in public housing and physical housing characteristics (log median number of rooms and log median year built) as exogenous variables. Standard errors are clustered at the tract level.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Table 2: The Welfare Effects of Public Housing Demolitions and Intermediate Counterfactuals

| Counterfactual scenario | Change from baseline (2010 Census) | | | | | |
| | Non-Hispanic White | | Black | | Hispanic | |
	Poor (1)	Non-poor (2)	Poor (3)	Non-poor (4)	Poor (5)	Non-poor (6)
Panel A. Results for renters						
Destroy buildings in tract	40	51	111	224	119	114
... and change neigborhood composition via demolitions	48	106	48	-80	59	35
... and change neighborhood composition via resorting	44	79	81	76	85	67
... and change housing prices (all channels for renters)	-46	-21	-134	-128	-132	-126
Panel B. Results for homeowners						
... and redistribute rents to homeowners (all channels for owners)	264	288	176	182	178	183
Panel C. Full equilibrium results						
Average welfare change across renters & owners	113	230	-75	39	-41	69
Aggregate welfare change across renters & owners (in millions)	$13.6	$180.9	-$10.7	$11.7	-$2.3	$17.5
Homeownership rate	51.3%	81.2%	19.0%	53.9%	29.4%	63.0%
Total households	120,840	786,279	142,858	300,190	57,177	254,458

Notes: This table reports the rent equivalent change in welfare for each counterfactual compared to a counterfactual with no public housing demolitions. A positive rent equivalent implies that households are better off in the indicated counterfactual relative to the counterfactual with no public housing demolitions. Panel A focuses on renter welfare. In the first row, we consider a counterfactual in which public housing is destroyed in each tract. In the second row, the Black and Hispanic population shares also adjust because of the removal of public housing residents. In the third row, these demographic variables further adjust as households re-optimize their location choices and displaced public housing residents seek market-based housing. The fourth row allows housing prices to adjust in addition. Panel B focuses on homeowner welfare when all endogenous outcomes adjust and the total change in rents in Chicago are redistributed as rental income. Panel C reports welfare results for renters and owners when all channels adjust to public housing demolitions. Statistics on total households by group in Cook County are based on the 2010 Census, and statistics on homeownership rates are from the 2008–2012 ACS.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Table 3: The Role of Spillovers in Generating City-Wide Rent Increases from Public Housing Demolitions

	Tracts with Demolitions (1)	Tracts without Demolitions (2)	All Tracts (3)
Number of tracts	57	1183	1240
Share of tracts	0.05	0.95	1.00
Average log rent increase	0.138	0.019	0.024
Share of total rent increase	0.26	0.74	1.00

Notes: This table describes the role of spillovers in generating increases in rents in Cook County after public housing demolitions. Columns 1 and 2 provide statistics for the groups of tracts that did and did not have public housing demolitions. Column 3 provide statistics for all tracts in Cook County. The third row reports the average log rent increase in a given group of tracts where the averages are weighted by the number of households living in each tract. The fourth row reports the share of the county-wide rent increase due to tracts with and without demolitions.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Online Appendices

A   Appendix Figures and Tables

Appendix Figure A.1: Changes in Housing Stock and Public Housing Demolitions, 2000–2010

(a) Share of Housing Units 0–10 Years Old

(b) Share of Housing Units 11-20 Years Old

(c) Share of Housing Units 21–30 Years Old

(d) Share of Housing Units 31+ Years Old

Notes: This figure plots the change in the share of housing units of the indicated age against the cumulative number of public housing units demolished from 2000–2010 as a share of the number of occupied housing units in 1990. We winsorize the public housing demolition share variable from above at 1 for 3 tracts. Each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.2: Tract-Level Rent Changes Due to Public Housing Demolitions Relative to Rents in Absence of Demolitions
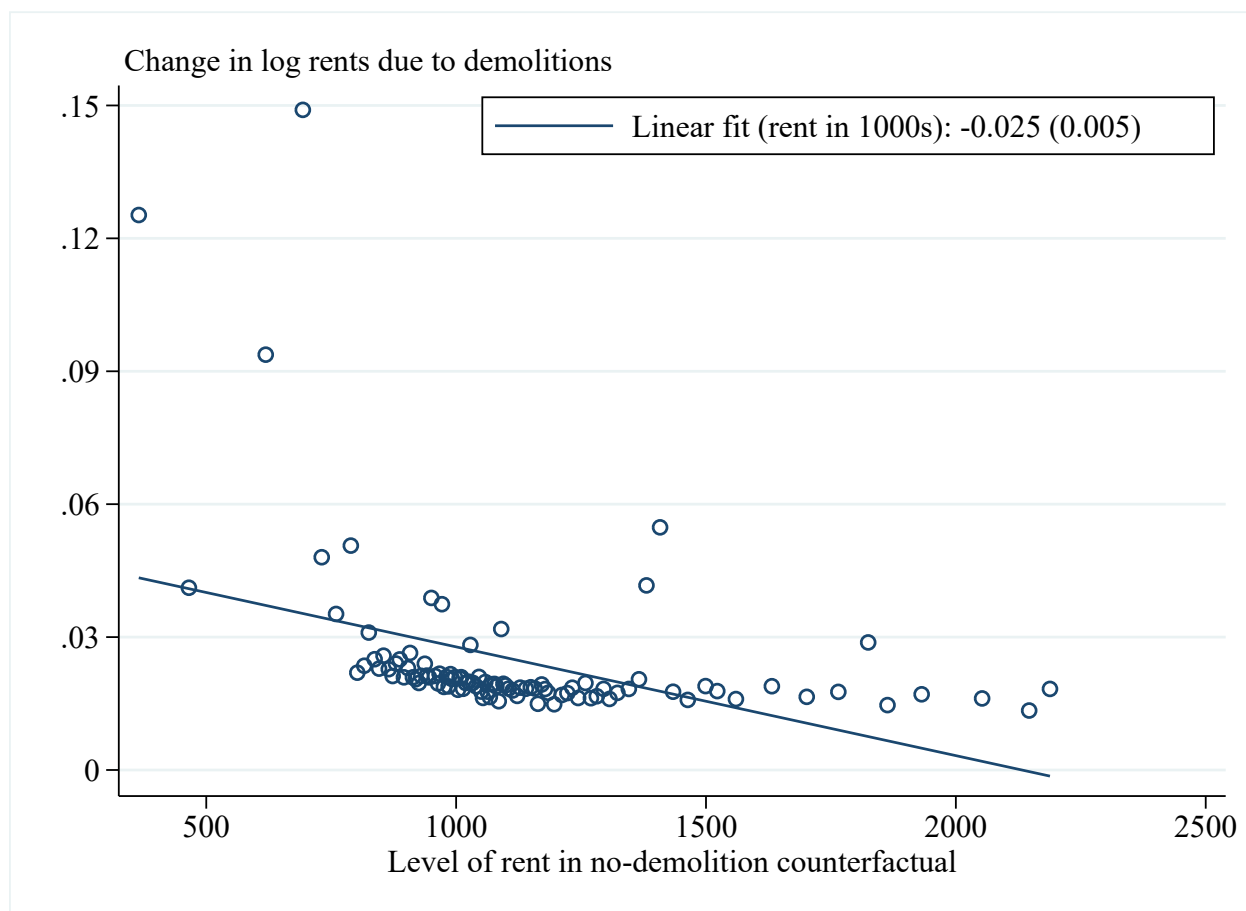
Change in log rents due to demolitions

Notes: This figure displays the change in median log rents due to public housing demolitions against the level of rents in the no-demolition counterfactual. We construct the dependent variable using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario where there are no demolitions. The linear fit expresses the relationship between the percent change and the level of rent in $1000's of dollars. The bin scatter is constructed for 100 percentiles.
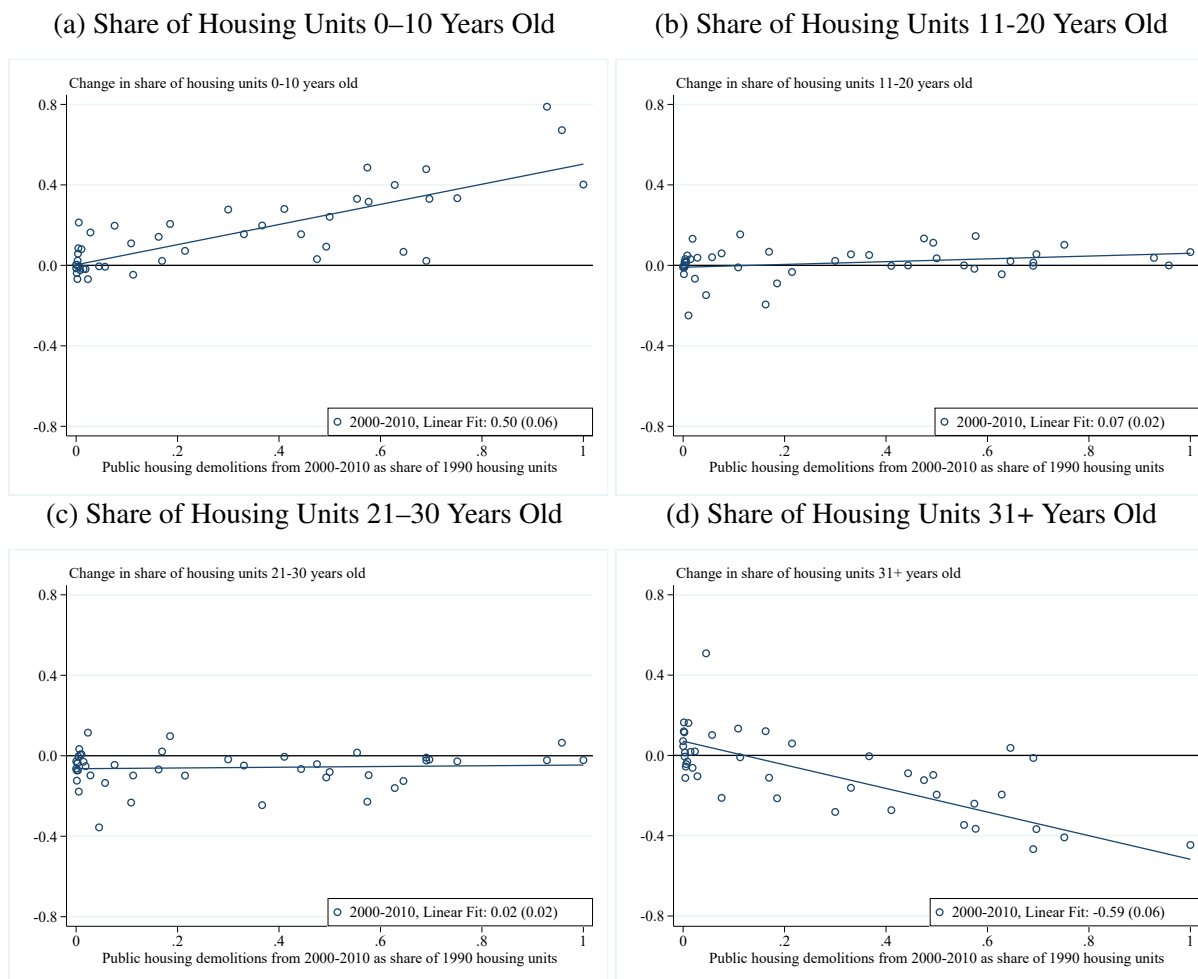
Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Appendix Figure A.3: Rents Increased by More in Non-Public-Housing Neighborhoods That Were Closer to Public Housing Demolitions
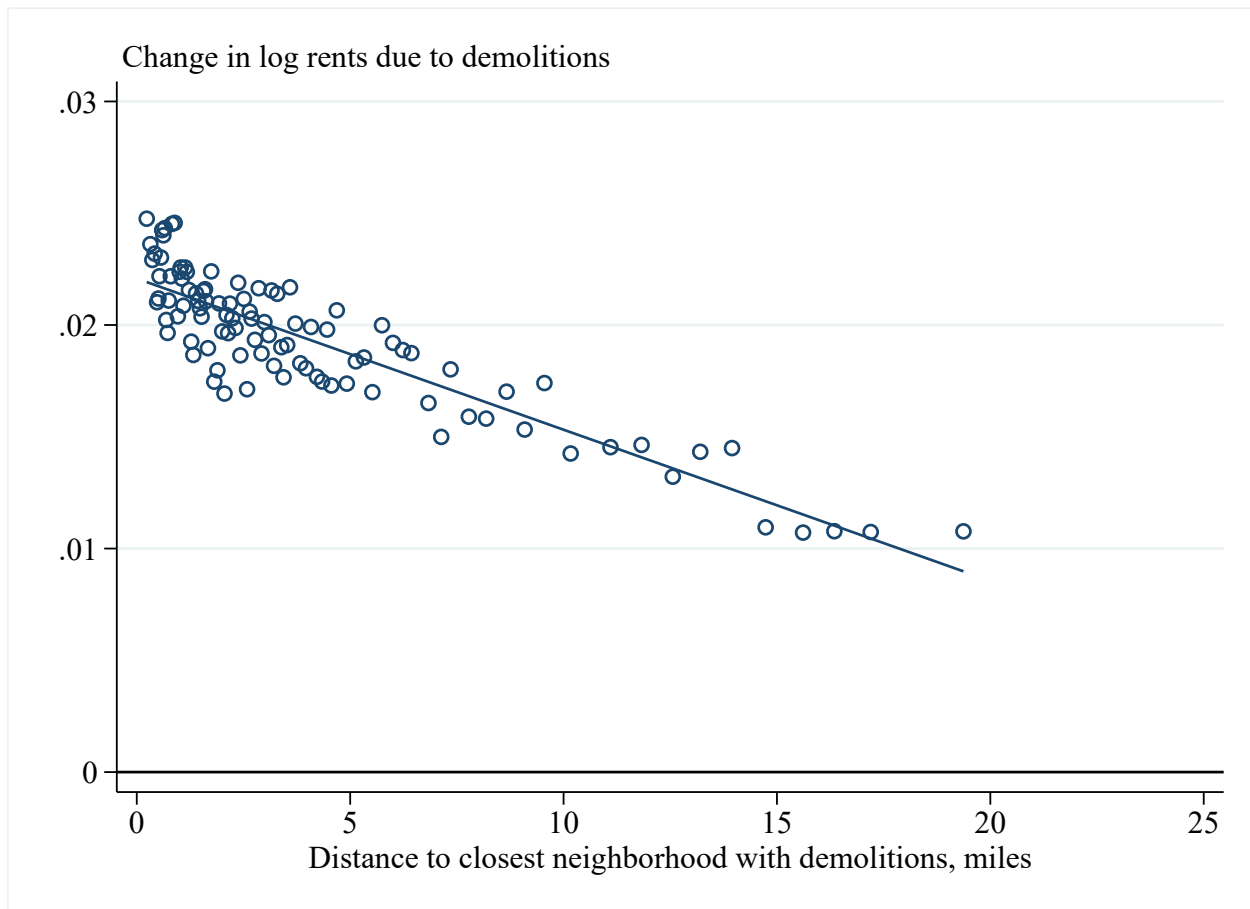
Change in log rents due to demolitions

Distance to closest neighborhood with demolitions, miles

Notes: Figure displays the change in median log rents due to public housing demolitions for tracts that did not have public housing against the distance to the closest tract with demolitions. We construct the dependent variable using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario where there are no demolitions. The bin scatter is constructed for 100 percentiles.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Appendix Figure A.4: Consequences of Public Housing Demolitions on Welfare Under Alternative Housing Supply Elasticities for Renters and Owners
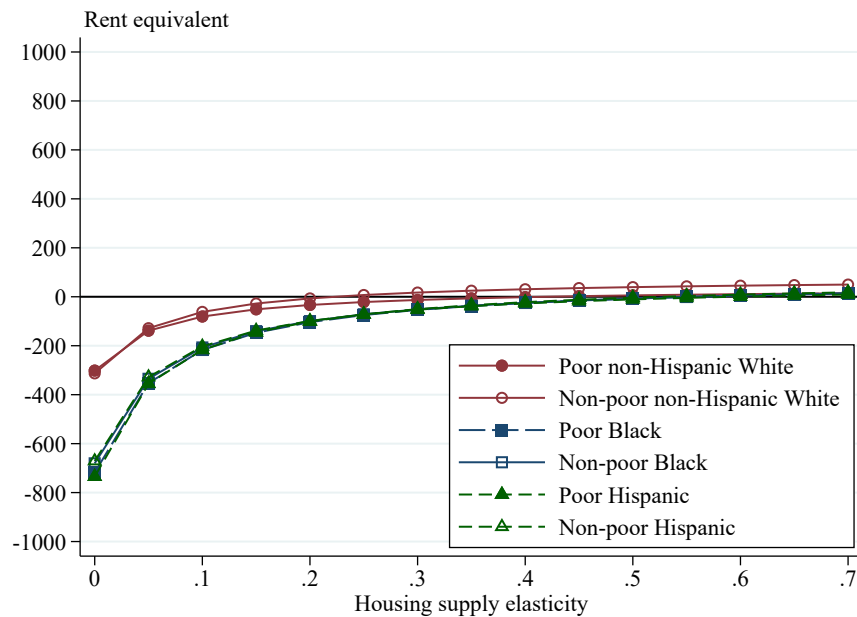
(a) Welfare Consequences for Renters

(b) Welfare Consequences for Owners
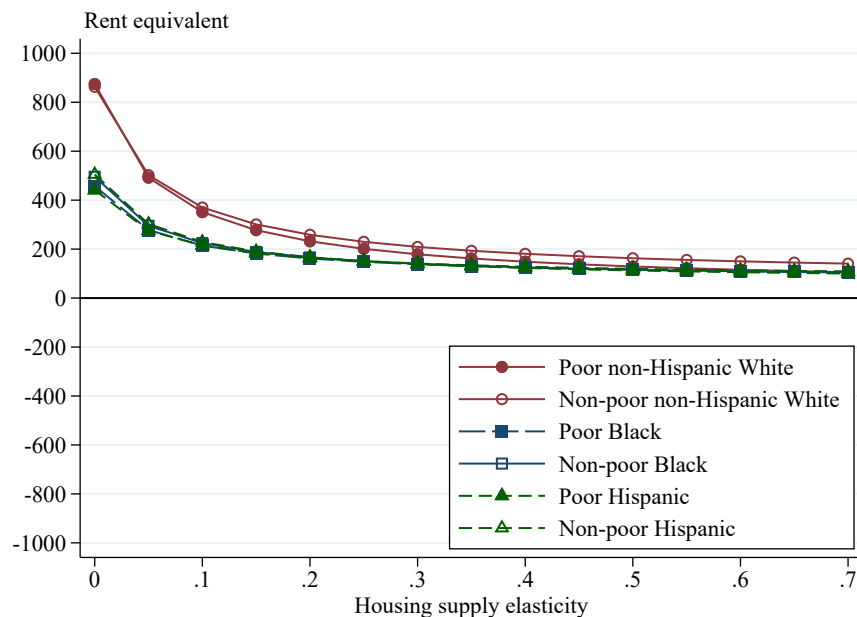
Notes: Figure displays outcomes from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Panel A shows the rent equivalent welfare effect of public housing demolitions for renters and Panel B shows analogous results for homeowners. Our baseline results are based on a housing supply elasticity of 0.163.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Appendix Figure A.5: Consequences of Public Housing Demolitions on Welfare Under Additional Redevelopment of Public Housing for Renters and Owners
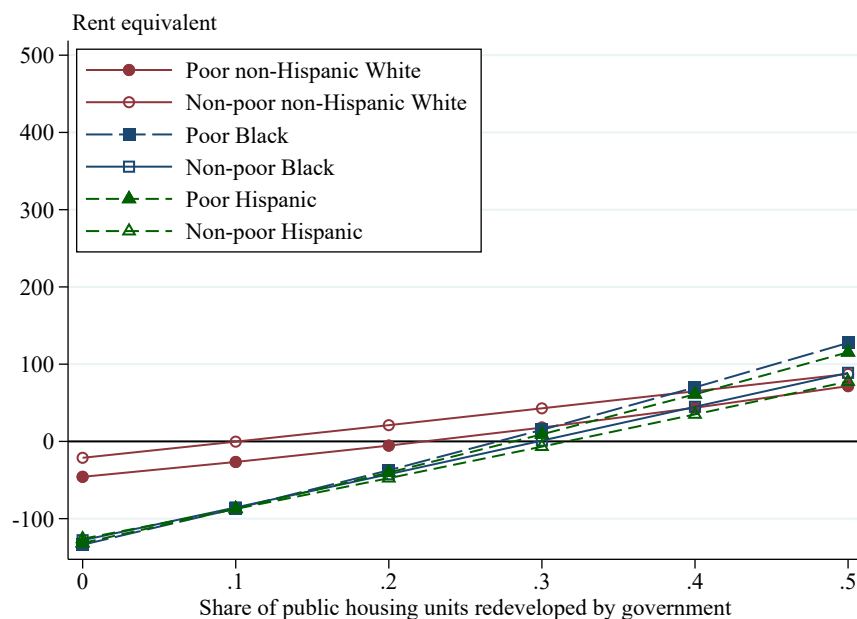
(a) Welfare Consequences for Renters

(b) Welfare Consequences for Owners
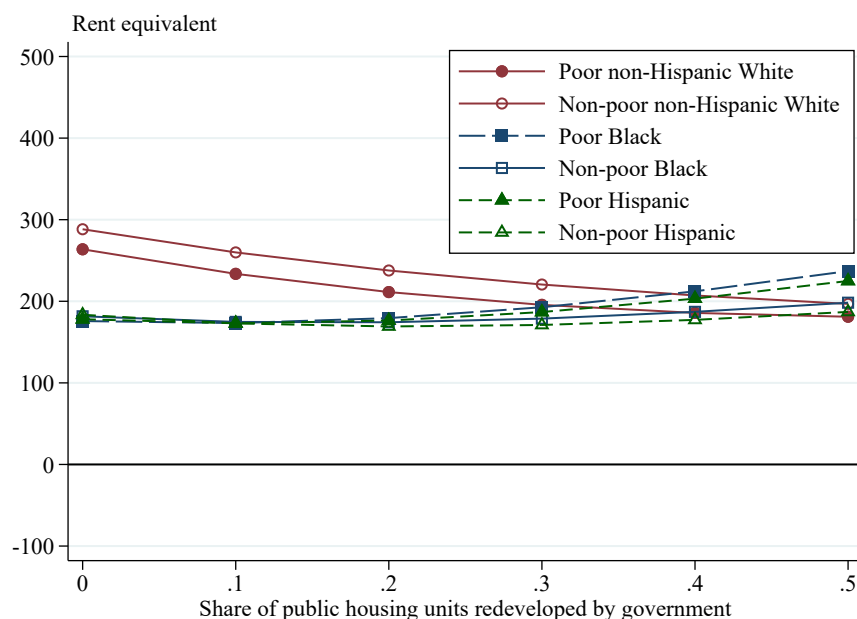
Notes: Figure displays outcomes from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. Panel A shows the rent equivalent welfare effect of public housing demolitions for renters and Panel B shows analogous results for homeowners. Our baseline results are given by 0 percent additional redevelopment.

Source: Authors' calculations using data from the Chicago Housing Authority and Census Bureau.

Appendix Table A.1: Land Use of Demolished Public Housing Units as of 2010

	Share	
Land use category	2010	2015
Vacant	0.38	0.35
Residential	0.40	0.43
Multi-Family	0.23	0.25
Single-Family Attached	0.14	0.16
Single-Family Detached	0.02	0.02
Commercial	0.08	0.08
Roadway or railroad	0.05	0.05
Institutional (school, government, and religious building)	0.04	0.04
Open Space (recreation)	0.04	0.04
Industrial	0.01	0.01
Under Construction	0.01	0.00

Notes: This table reports the share of demolished public housing units with the indicated land use category as of 2010 and 2015.

Source: Authors' calculations using data from the Chicago Housing Authority and Chicago Metropolitan Agency for Planning Land Use Inventory.

Appendix Table A.2: OLS Estimates of Neighborhood Preference Parameters

	Preference parameters for indicated group		
	Non-Hispanic White (1)	Black (2)	Hispanic (3)
Panel A: Poor Households			
Log median rent	-0.0455***	-0.0286***	-0.0348***
	(0.00581)	(0.00529)	(0.00858)
Black population share	-0.0969***	0.239***	0.272***
	(0.0164)	(0.0202)	(0.0253)
Hispanic population share	-0.0676***	0.0465***	0.340***
	(0.0138)	(0.0133)	(0.0224)
Log median household income	0.00513	-0.0110**	-0.0321***
	(0.00392)	(0.00544)	(0.00762)
Public housing units as a share of housing stock	-0.0120	-0.00338	-0.0380
	(0.0140)	(0.0216)	(0.0279)
Panel B: Non-poor Households			
Log median rent	-0.00175	-0.00625*	-0.0218***
	(0.00234)	(0.00333)	(0.00706)
Black population share	-0.120***	0.208***	0.174***
	(0.00958)	(0.0177)	(0.0203)
Hispanic population share	-0.138***	0.0799***	0.265***
	(0.00580)	(0.00812)	(0.0180)
Log median household income	0.0111***	0.00366	-0.0200***
	(0.00225)	(0.00327)	(0.00629)
Public housing units as a share of housing stock	-0.00653	-0.0497***	-0.0374*
	(0.00894)	(0.0137)	(0.0209)
Specifications include:			
Year Fixed Effects	✓	✓	✓
Tract Fixed Effects	✓	✓	✓
Log median number of rooms	✓	✓	✓
Log median year built	✓	✓	✓
Homeownership share	✓	✓	✓
Land use variables	✓	✓	✓
Observations (tract-by-year)	2,480	2,480	2,480
Number of tracts	1,240	1,240	1,240

Notes: This table presents regression results of preference parameters for a static logit location choice model using household counts across census tracts in Cook County for 2000 and 2010. We estimate preference parameters separately by race/ethnicity and income group. Poor households have income below $20,000, and non-poor households have income above $20,000. These estimates do not use our preferred instrumental variable approach. Standard errors are clustered at the tract level.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Table A.3: Instrumental Variable Estimates of Neighborhood Preference Parameters, Poor Non-Hispanic White and Black Households, Robustness

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	IV rings: 2–3, 3–5 miles (4)	IV rings: 2–3, 3–5, 5–10 miles (5)	Add 1990 & 1990-2000 controls (6)	Drop < 1 mile Cabrini-Green (7)
Poor Non-Hispanic White Households							
Log median rent	-0.455***	-0.430***	-0.458***	-0.493***	-0.469***	-0.620***	-0.487***
	(0.0552)	(0.0667)	(0.0558)	(0.0538)	(0.0543)	(0.148)	(0.0602)
Black population share	-0.165*	-0.134	-0.158*	-0.217**	-0.177**	-0.0730	-0.134
	(0.0929)	(0.0873)	(0.0949)	(0.0992)	(0.0886)	(0.0973)	(0.101)
Hispanic population share	-0.132***	-0.0988**	-0.132***	-0.118***	-0.0879**	-0.144***	-0.138***
	(0.0365)	(0.0388)	(0.0366)	(0.0407)	(0.0384)	(0.0478)	(0.0394)
Log median household income	0.0886***	0.0864***	0.0894***	0.0953***	0.106***	0.157***	0.0832***
	(0.0232)	(0.0255)	(0.0234)	(0.0265)	(0.0251)	(0.0411)	(0.0241)
PH units as a share of housing stock	-0.450***	-0.409***	-0.444***	-0.480***	-0.451***	-0.409***	-0.494***
	(0.107)	(0.112)	(0.108)	(0.110)	(0.106)	(0.136)	(0.122)
Poor Black Households							
Log median rent	-0.241***	-0.106***	-0.243***	-0.256***	-0.201***	-0.298***	-0.273***
	(0.0333)	(0.0202)	(0.0337)	(0.0294)	(0.0264)	(0.0791)	(0.0359)
Black population share	0.265***	0.176***	0.267***	0.234***	0.261***	0.245***	0.299***
	(0.0542)	(0.0296)	(0.0553)	(0.0562)	(0.0438)	(0.0563)	(0.0581)
Hispanic population share	0.00195	0.0180	0.00184	0.0450*	0.0245	-0.0684***	-0.00559
	(0.0216)	(0.0137)	(0.0217)	(0.0236)	(0.0205)	(0.0256)	(0.0241)
Log median household income	0.0353***	0.0114	0.0357***	0.0503***	0.0273**	0.0753***	0.0346**
	(0.0127)	(0.00693)	(0.0128)	(0.0149)	(0.0120)	(0.0218)	(0.0137)
PH units as a share of housing stock	-0.242***	-0.121***	-0.242***	-0.242***	-0.196***	-0.240***	-0.266***
	(0.0639)	(0.0341)	(0.0639)	(0.0633)	(0.0545)	(0.0725)	(0.0769)

Notes: This table presents regression results of preference parameters for a static logit location choice model using household counts across census tracts in Cook County for 2000 and 2010. We estimate preference parameters separately by race/ethnicity and income group. Poor households have income below $20,000, and non-poor households have income above $20,000. Log median rent, Black and Hispanic population share, and log median income are instrumented following Bayer et al. (2007), where we take changes in public housing and physical housing characteristics (median number of rooms and median year built) as exogenous variables. Column 1 reports results from our baseline specification (also reported in Table 1). The instrumental variables in this specification are based on rings that are 3–5, 5–10, and 10–20 miles away. Column 2 adds separate control variables for averages of the median room, median year built, and public housing share variables in tracts that are 0–1, 1–2, and 2–3 miles away. Column 3 adds the homicide rate as a control. Columns 4 and 5 use the baseline covariates and instrumental variables based on rings that are 2–3 and 3–5 miles away or 2–3, 3–5, and 5–10 miles away. Column 6 adds interactions between fixed effects for year and the 1990 level of log median rent, log median household income, and share of residents with a college education, along with interactions between fixed effects for year and changes from 1990 to 2000 in these three variables. Column 7 drops tracts that are within 1 mile of the Cabrini-Green Homes. Standard errors are clustered at the tract level.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Table A.4: Instrumental Variable Estimates of Neighborhood Preference Parameters, Poor Hispanic and Other Race/Ethnicity Households, Robustness

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	IV rings: 2–3, 3–5 miles (4)	IV rings: 2–3, 3–5, 5–10 miles (5)	Add 1990 & 1990-2000 controls (6)	Drop < 1 mile Cabrini-Green (7)
Poor Hispanic Households							
Log median rent	-0.246***	-0.136***	-0.243***	-0.243***	-0.220***	-0.208***	-0.274***
	(0.0396)	(0.0328)	(0.0394)	(0.0343)	(0.0343)	(0.0714)	(0.0415)
Black population share	0.273***	0.209***	0.274***	0.284***	0.273***	0.172***	0.303***
	(0.0573)	(0.0405)	(0.0574)	(0.0553)	(0.0481)	(0.0465)	(0.0595)
Hispanic population share	0.307***	0.317***	0.307***	0.351***	0.320***	0.160***	0.300***
	(0.0283)	(0.0240)	(0.0281)	(0.0293)	(0.0286)	(0.0265)	(0.0301)
Log median household income	0.0128	-0.00269	0.0122	0.0262	0.00915	0.0383*	0.00945
	(0.0151)	(0.0119)	(0.0149)	(0.0162)	(0.0153)	(0.0203)	(0.0158)
PH units as a share of housing stock	-0.270***	-0.164***	-0.269***	-0.257***	-0.240***	-0.223***	-0.296***
	(0.0733)	(0.0583)	(0.0722)	(0.0700)	(0.0668)	(0.0601)	(0.0852)
Poor Other Race/Ethnicity Households							
Log median rent	0.0832**	0.00218	0.0877***	0.0338	-0.00539	0.530***	0.106***
	(0.0334)	(0.0273)	(0.0338)	(0.0319)	(0.0308)	(0.152)	(0.0367)
Black population share	-0.182***	-0.000359	-0.180***	-0.179***	-0.198***	-0.264***	-0.196***
	(0.0363)	(0.0237)	(0.0365)	(0.0338)	(0.0336)	(0.0909)	(0.0382)
Hispanic population share	0.0686***	0.180***	0.0689***	0.0274	0.0551**	0.0931*	0.0745***
	(0.0249)	(0.0214)	(0.0250)	(0.0237)	(0.0237)	(0.0506)	(0.0257)
Log median household income	-0.0322***	-0.0186**	-0.0331***	-0.0386***	-0.00608	-0.152***	-0.0359***
	(0.0116)	(0.00849)	(0.0117)	(0.0121)	(0.0119)	(0.0424)	(0.0118)
PH units as a share of housing stock	-0.0293	-0.0297	-0.0264	-0.0967**	-0.119***	0.199	-0.0218
	(0.0497)	(0.0385)	(0.0500)	(0.0470)	(0.0457)	(0.131)	(0.0566)

Notes: See notes to Appendix Table A.3.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Table A.5: Instrumental Variable Estimates of Neighborhood Preference Parameters, Non-Poor Non-Hispanic White and Black Households, Robustness

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	IV rings: 2–3, 3–5 miles (4)	IV rings: 2–3, 3–5, 5–10 miles (5)	Add 1990 & 1990-2000 controls (6)	Drop < 1 mile Cabrini-Green (7)
Non-Poor Non-Hispanic White Households							
Log median rent	-0.0564***	-0.0627***	-0.0582***	-0.0741***	-0.0727***	-0.0739***	-0.0575***
	(0.00908)	(0.0112)	(0.00925)	(0.0108)	(0.0109)	(0.0269)	(0.00992)
Black population share	-0.134***	-0.0988***	-0.132***	-0.155***	-0.139***	-0.0856***	-0.133***
	(0.0163)	(0.0146)	(0.0167)	(0.0191)	(0.0179)	(0.0139)	(0.0178)
Hispanic population share	-0.142***	-0.115***	-0.142***	-0.158***	-0.145***	-0.102***	-0.141***
	(0.00712)	(0.00792)	(0.00718)	(0.00873)	(0.00830)	(0.00794)	(0.00727)
Log median household income	0.0217***	0.0194***	0.0221***	0.0170***	0.0246***	0.0242***	0.0211***
	(0.00387)	(0.00420)	(0.00397)	(0.00480)	(0.00480)	(0.00743)	(0.00383)
PH units as a share of housing stock	-0.0639***	-0.0537***	-0.0628***	-0.0849***	-0.0812***	-0.0502***	-0.0697***
	(0.0153)	(0.0184)	(0.0158)	(0.0183)	(0.0179)	(0.0186)	(0.0154)
Non-Poor Black Households							
Log median rent	-0.0368***	0.0299**	-0.0378***	-0.0508***	-0.0275**	0.0893***	-0.0465***
	(0.0109)	(0.0119)	(0.0110)	(0.0115)	(0.0113)	(0.0331)	(0.0117)
Black population share	0.220***	0.225***	0.222***	0.211***	0.224***	0.177***	0.228***
	(0.0212)	(0.0178)	(0.0215)	(0.0227)	(0.0204)	(0.0218)	(0.0228)
Hispanic population share	0.0872***	0.134***	0.0872***	0.0832***	0.0677***	0.0571***	0.0844***
	(0.00876)	(0.0107)	(0.00870)	(0.00907)	(0.00888)	(0.0122)	(0.00909)
Log median household income	0.00978**	-0.00748	0.0101**	0.00748	-0.00115	-0.0124	0.0101**
	(0.00389)	(0.00477)	(0.00393)	(0.00480)	(0.00489)	(0.00951)	(0.00397)
PH units as a share of housing stock	-0.0848***	0.00536	-0.0829***	-0.102***	-0.0824***	-0.0204	-0.0920***
	(0.0201)	(0.0185)	(0.0201)	(0.0219)	(0.0203)	(0.0253)	(0.0225)

Notes: See notes to Appendix Table A.3.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Table A.6: Instrumental Variable Estimates of Neighborhood Preference Parameters, Non-Poor Hispanic and Other Race/Ethnicity Households, Robustness

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	IV rings: 2–3, 3–5 miles (4)	IV rings: 2–3, 3–5, 5–10 miles (5)	Add 1990 & 1990-2000 controls (6)	Drop < 1 mile Cabrini-Green (7)
Non-Poor Hispanic Households							
Log median rent	-0.103***	-0.0213	-0.0994***	-0.111***	-0.107***	0.0818	-0.115***
	(0.0263)	(0.0231)	(0.0261)	(0.0253)	(0.0253)	(0.0606)	(0.0277)
Black population share	0.178***	0.152***	0.177***	0.195***	0.172***	0.0546**	0.189***
	(0.0299)	(0.0231)	(0.0296)	(0.0308)	(0.0284)	(0.0273)	(0.0306)
Hispanic population share	0.261***	0.296***	0.261***	0.282***	0.253***	0.152***	0.257***
	(0.0199)	(0.0195)	(0.0198)	(0.0204)	(0.0203)	(0.0205)	(0.0205)
Log median household income	-0.00313	-0.0170**	-0.00392	0.00508	-0.00362	-0.0273*	-0.00521
	(0.00916)	(0.00789)	(0.00905)	(0.0101)	(0.0103)	(0.0165)	(0.00932)
PH units as a share of housing stock	-0.127***	-0.0330	-0.126***	-0.134***	-0.132***	-0.0386	-0.139***
	(0.0415)	(0.0364)	(0.0407)	(0.0422)	(0.0417)	(0.0377)	(0.0468)
Non-Poor Other Race/Ethnicity Households							
Log median rent	0.0978***	0.0738***	0.0995***	0.0427*	0.0249	0.526***	0.119***
	(0.0242)	(0.0199)	(0.0245)	(0.0219)	(0.0207)	(0.136)	(0.0265)
Black population share	-0.174***	-0.0440**	-0.173***	-0.190***	-0.195***	-0.221***	-0.179***
	(0.0275)	(0.0193)	(0.0277)	(0.0237)	(0.0231)	(0.0834)	(0.0290)
Hispanic population share	-0.0497***	0.0676***	-0.0496***	-0.0926***	-0.0774***	0.0225	-0.0432**
	(0.0191)	(0.0177)	(0.0192)	(0.0169)	(0.0171)	(0.0466)	(0.0199)
Log median household income	-0.0114	-0.0141**	-0.0117	-0.0195**	0.00207	-0.131***	-0.0127
	(0.00870)	(0.00696)	(0.00875)	(0.00827)	(0.00788)	(0.0381)	(0.00910)
PH units as a share of housing stock	0.0285	0.0653**	0.0298	-0.0429	-0.0496*	0.227*	0.0465
	(0.0369)	(0.0281)	(0.0369)	(0.0294)	(0.0272)	(0.123)	(0.0432)

Notes: See notes to Appendix Table A.3.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Table A.7: Instrumental Variable Estimates of Neighborhood Preference Parameters, Heterogeneity by 1990 Median Income

	Poor Households				Non-Poor Households			
	Non-Hisp. White	Black	Hispanic	Other	Non-Hisp. White	Black	Hispanic	Other
Log median rent	-0.456***	-0.239***	-0.244***	0.0813**	-0.0573***	-0.0355***	-0.102***	0.0959***
	(0.0551)	(0.0332)	(0.0394)	(0.0335)	(0.00921)	(0.0109)	(0.0263)	(0.0242)
Black population share	-0.176*	0.269***	0.274***	-0.189***	-0.137***	0.222***	0.178***	-0.177***
	(0.0949)	(0.0552)	(0.0584)	(0.0368)	(0.0167)	(0.0215)	(0.0304)	(0.0279)
Hispanic population share	-0.133***	0.00231	0.307***	0.0678***	-0.142***	0.0875***	0.261***	-0.0501***
	(0.0367)	(0.0215)	(0.0282)	(0.0249)	(0.00719)	(0.00877)	(0.0199)	(0.0190)
Log median household income	0.0864***	0.0354***	0.0122	-0.0332***	0.0213***	0.00989**	-0.00351	-0.0114
	(0.0230)	(0.0125)	(0.0150)	(0.0115)	(0.00384)	(0.00385)	(0.00915)	(0.00860)
PH units ... × Decile 1 of 1990 median HH inc.	-0.479***	-0.235***	-0.273***	-0.0480	-0.0716***	-0.0793***	-0.129***	0.0221
	(0.113)	(0.0671)	(0.0784)	(0.0524)	(0.0147)	(0.0204)	(0.0443)	(0.0384)
PH units ... × Deciles 2-10 of 1990 median HH inc.	-0.253*	-0.276***	-0.243***	0.0867	-0.0170	-0.114***	-0.111**	0.0592
	(0.133)	(0.0871)	(0.0877)	(0.0630)	(0.0304)	(0.0290)	(0.0463)	(0.0555)

Notes: This table presents regression results of preference parameters for a static logit location choice model using household counts across census tracts in Cook County for 2000 and 2010. We estimate preference parameters separately by race/ethnicity and income group. Poor households have income below $20,000, and non-poor households have income above $20,000. Log median rent, Black and Hispanic population share, and log median income are instrumented following Bayer et al. (2007), where we take changes in public housing and physical housing characteristics (median number of rooms and median year built) as exogenous variables. We allow the coefficients on the public housing share variables to differ based on whether the tract's 1990 median household income is in the first decile in Cook County (which accounts for 10% of tracts and 90% of demolitions) or in deciles 2–10. Standard errors are clustered at the tract level.

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

B Details on Equilibrium Solver

Given exogenous location characteristics $(\mathbf{x}, \xi^{\mathbf{k}})$ and preference parameters $(\alpha^k)_{k=1}^K$, we want to find a vector of prices and endogenous amenities $(\mathbf{p}, \mathbf{b}, \mathbf{h})$ that solves simultaneously the following system of equations:

$$\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha) = \mathcal{S}_j(p_j) \qquad \forall j = 1, ..., J \qquad (\text{B.1})$$

$$\frac{\mathcal{D}_j^B(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha)}{\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha)} = b_j \qquad \forall j = 1, ..., J \qquad (\text{B.2})$$

$$\frac{\mathcal{D}_j^H(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha))}{\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha))} = h_j \qquad \forall j = 1, ..., J. \qquad (\text{B.3})$$

In what follows, we describe our algorithm solver. Because $(\mathbf{x}, \xi^{\mathbf{k}})$ and $(\alpha^k)_{k=1}^K$ are fixed, we suppress them to simplify notation.

The first step is to construct an excess demand function, for both housing and demographic composition. Those are given as follows:

$$\mathcal{EDH}(\mathbf{p}, \mathbf{b}, \mathbf{h}) = \begin{bmatrix} \mathcal{D}_1(\mathbf{p}, \mathbf{b}, \mathbf{h}) - \mathcal{S}_1(p_1) \\ \vdots \\ \mathcal{D}_J(\mathbf{p}, \mathbf{b}, \mathbf{h}) - \mathcal{S}_J(p_J) \end{bmatrix} \qquad (\text{B.4})$$

$$\mathcal{EDD}(\mathbf{p}, \mathbf{b}, \mathbf{h}) = \begin{bmatrix} \frac{\mathcal{D}_1^B(\mathbf{p,b,h})}{\mathcal{D}_1(\mathbf{p,b,h})} - b_1 \\ \vdots \\ \frac{\mathcal{D}_J^B(\mathbf{p,b,h})}{\mathcal{D}_J(\mathbf{p,b,h})} - b_J \\ \frac{\mathcal{D}_1^H(\mathbf{p,b,h})}{\mathcal{D}_1(\mathbf{p,b,h})} - h_1 \\ \vdots \\ \frac{\mathcal{D}_J^H(\mathbf{p,b,h})}{\mathcal{D}_J(\mathbf{p,b,h})} - h_J. \end{bmatrix} \qquad (\text{B.5})$$

Observe that an equilibrium is defined whenever $\mathcal{EDH}(\mathbf{p}, \mathbf{b}, \mathbf{h}) = 0$ and $\mathcal{EDD}(\mathbf{p}, \mathbf{b}, \mathbf{h}) = 0$. To find the zeroes of such a system of equations, we set an initial guess $(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0)$ and follow an iterative algorithm described as follows:

1. For a given guess $(\mathbf{p}^n, \mathbf{b}^n, \mathbf{h}^n)$, evaluate excess demand functions and obtain values $\mathcal{EDH}^n$ and $\mathcal{EDD}^n$.

2. Update the guess as follows:

   - $\mathbf{p}^{n+1} = \mathbf{p}^n + \tau \cdot \mathcal{EDH}^n$

   - $\begin{bmatrix} \mathbf{b}^{n+1} \\ \mathbf{h}^{n+1} \end{bmatrix} = \begin{bmatrix} \mathbf{b}^n \\ \mathbf{h}^n \end{bmatrix} - \tau \cdot \mathcal{EDD}^n.$

The update on prices and demographic composition go in opposite directions because prices act as a congestion force in our model whereas demographic act as an agglomeration force.

The tuning parameter τ is fixed by the practitioner. Higher values of τ lead to a faster but more unstable fixed-point search. In our application we set $\tau = 0.2$ and our initial value equal to the observed equilibrium in the data. We set our tolerance criterion as follows:

$$\max\left\{||\mathcal{EDH}(\mathbf{p}^n, \mathbf{b}^n, \mathbf{h}^n)||_\infty, ||\mathcal{EDD}(\mathbf{p}^n, \mathbf{b}^n, \mathbf{h}^n)||_\infty\right\} < e^{-10}.$$

A fixed point of the system of equations (B.1)–(B.3) can also be found using a non-linear optimization package. In that case, we define our objective function as follows:

$$\left(\sum_j \mathcal{EDH}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}) + \sum_j \mathcal{EDD}_j(\mathbf{p}, \mathbf{b}, \mathbf{h})\right)^2$$

To minimize the previous function, we use the optimization algorithm L-BFGS, which is part of the package *Optim* in Julia. We use the *Accelerated Gradient Descent* algorithm with *automatic differentiation* given by *forward differences*.

Both methods deliver the same answer, but due to the large dimension of the solution space $(3 \cdot 1230 = 3714)$, the iterative algorithm is orders of magnitude faster and finds a solution in minutes or seconds.