Neighborhood Revitalization and Inequality: Evidence from Chicago's Public Housing Demolitions*

Milena Almagro
University of Chicago and NBER

Eric Chyn
UT-Austin and NBER

Bryan A. Stuart
Federal Reserve Bank of Philadelphia

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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 minority households, especially those who were displaced from public housing, and increased for higher-income 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, I31.

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*[*] **Almagro:** University of Chicago, Booth School of Business, 5807 S Woodlawn Ave, Chicago, IL 60637, Email: milena.almagro@chicagobooth.edu. **Chyn:** Department of Economics, University of Texas at Austin, 2225 Speedway, Stop C3100, Austin, TX 78712 and the National Bureau of Economic Research (NBER), Email: eric.chyn@austin.utexas.edu. **Stuart:** Research Department, Federal Reserve Bank of Philadelphia, 10 Independence Mall, Philadelphia, PA 19106 and IZA, Email: bryan.stuart@phil.frb.org. The views expressed in this paper are solely those of the authors and do not necessarily reflect the views of the Federal Reserve Bank of Philadelphia or the Federal Reserve System. We thank Francisco Gallego, Jesse Gregory, and Andrii Parkhomenko for insightful discussions. We are also grateful for helpful comments from Dionissi Aliprantis, Donald Davis, Tomas Dominguez-Iino, Fernando Ferreira, Cecile Gaubert, Elisa Giannone, Edward Glaeser, Jessie Handbury, Anders Humlum, Erik Hurst, Jeff Lin, Matt Notowidigdo, Esteban Rossi-Hansberg, Maisy Wong, and seminar participants at the University of Wisconsin, Notre Dame Spatial Workshop, OIGI Fall Conference, Atlanta Fed, University of Rochester, SMU, Central Bank of Colombia, University of Chicago Booth School of Business, Barcelona Summer Forum Trade Workshop, the NBER Real Estate and Urban Economics Summer Institute 2022 Workshop, the Princeton Spatial Conference, the Federal Reserve Bank of Cleveland, the Urban Economics Association Fall 2022 Meeting, the 2022 Inaugural Conference on Economic Opportunity, and the University of California, Berkeley. Elias Van Emmerick and Vanessa Ntungwanayo provided outstanding research assistance. Any errors are our own.

1 Introduction

Concerns over inequality and the concentration of poverty within cities have prompted governments 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 neighborhood redevelopment 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 places. Assessing the welfare consequences of these local investment 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 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. Public housing demolitions occurred in 5 percent of the tracts in Cook County.

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 large increases in the White population share alongside decreases in the share of Black and Hispanic residents. 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 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, individuals may explicitly care about the presence of low-income housing in their neighborhood (Diamond, McQuade and Qian, 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 re-sorting 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 focus on distant neighborhoods and use the changes in their housing market characteristics and the reductions in their public housing due to demolitions 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, Levinsohn and Pakes (1995) and Bayer, Ferreira and McMillan (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 (2024).

Our estimates of the residential choice model quantify the tradeoffs that households make when deciding where to live. We find that households prefer lower rents, a higher share of residents of their own race/ethnicity, and higher-income residents. The results are broadly in line with findings from prior studies such as Galiani, Murphy and Pantano (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 \$122 more in annual rents for a 1 percentage point reduction in the share of public housing in their neighborhood.

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 \$285 (2 percent) increase in annual rent equivalent units. In contrast, minority households generally saw declines in welfare with negative impacts of $-\$247$ for poor Black households, $-\$429$ for non-poor Black households, and sizable but smaller decreases in welfare for Hispanic households. Since White households constitute most of the population, there is a slight overall \$16 gain in average rent equivalent welfare when we aggregate the welfare of all non-Hispanic White, Black, and Hispanic households. The racial disparities in the welfare impacts of demolition remain when we extend our model to account for moving costs.

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

[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, Murphy and Pantano (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 \$134 per year to have a 1 percentage point increase in the share of their own race neighbors.

renters from all demographic groups are made better off when public housing buildings are demolished and nothing else changes. However, once accounting for a variety of equilibrium responses, nearly all demographic groups are made worse off from demolitions due to large increases in housing prices. Welfare losses for White renters are smaller 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 nearly 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. Non-poor White households are significantly better off due to their relatively high rates of home ownership, while poor White households are nearly indifferent. In contrast, Black and Hispanic households are worse off.

In addition to studying overall impacts, we use our estimated structural model to quantify spatial spillovers in the housing market. While demolitions on average increase rents by 2.5 percent across all census tracts in Cook County, the magnitude of the effects sharply differs based on whether a neighborhood was directly targeted. Neighborhoods with a public housing demolition saw rental prices increase by 15.6 percent on average, while areas without a demolition saw rents go up by 2.0 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.

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. These results highlight the potential for housing policy to shape and influence equity considerations associated with urban renewal programs.

As a final analysis, we combine an augmented version of our spatial model with novel administrative data to study welfare for residents of public housing. Unlike the households studied in our main analysis, those in public housing were directly affected by the demolitions because they were *forced* to relocate to new housing units using rental vouchers. Our analysis relies on individual-level panel data that contains information on public housing resident locations before and after building demolitions began in Chicago. Our empirical strategy compares residents of buildings destroyed during the initial wave of demolitions during the 1990s to their neighbors living in nearby non-demolished public housing. This latter group of households constitutes our control group as they did not have access to vouchers and their buildings remained untouched for several years. This setting creates the conditions for a rare natural experiment that allows us

[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 30 percent, lower-income minority households avoid welfare losses.

to obtain credible estimates of neighborhood preferences and moving costs.[4]

Our main finding for public housing residents is that displacement from public housing led to large welfare losses that were driven by displacement costs.[5] We find that the provision of housing vouchers generates substantively large welfare gains by increasing neighborhood choice, although the benefits are not sufficient to offset entirely the costs of relocation. We conclude by combining welfare estimates for the broader population in Cook County—estimated using census data—and those displaced from public housing. Our preferred estimates for this exercise show that the large welfare losses experienced by minority households and those removed from public housing outweigh the welfare gains experienced by non-poor White households.

2 Related Literature

The main contribution of this paper is to study the welfare and distributional impacts of one of the largest place-based investment programs in U.S. history using a spatial equilibrium model. Our approach captures equilibrium responses in both housing prices and demographic composition, while maintaining computational tractability and being estimable with commonly available aggregate data. Methodologically, this paper is most closely related to prior empirical studies that model endogenous amenities to examine sorting and welfare within cities (e.g., Su, 2022; De la Roca, Parkhomenko and Velásquez-Cabrera, 2023; Almagro and Domínguez-Iino, 2024; Couture et al., 2024; Redding and Sturm, 2024). We build on previous work by estimating a model that features substantially more preference heterogeneity in terms of income and race, which is key to our study of sorting in the racially segregated environment of Chicago.

Our analysis also relates to the literature that quantifies the forces that drive neighborhood sorting in the U.S. (Bayer, Ferreira and McMillan, 2007; Bayer et al., 2016; Caetano and Maheshri, 2021). Relative to this body of work, we contribute by studying an equilibrium framework with a housing market that explicitly accounts for housing supply responses. The supply responses are central to our counterfactual analysis that sheds light on how alternative urban redevelopment policies may shape racial disparities.

In addition, this paper topically connects to a set of reduced-form studies estimating the impacts of urban renewal and public housing demolition in the U.S. (Collins and Shester, 2013; Aliprantis and Hartley, 2015; Sandler, 2017; Tach and Emory, 2017; Blanco, 2023; Staiger, Palloni and Voorheis, 2024). These studies use difference-in-difference and related approaches to provide compelling evidence that HOPE VI and similar redevelopment programs can have substantively large impacts on neighborhood demographic composition. Our analysis builds on these findings by being the first to quantify the effects of large-scale urban redevelopment programs on welfare using a spatial equilibrium framework with heterogeneous agents. The structural framework delivers our key finding that the HOPE VI program had heterogeneous welfare impacts and appears to have worsened racial disparities in our setting.

Our final analysis of public housing residents displaced by demolitions contributes to an emerging literature using structural approaches to study neighborhood mobility for families receiving housing assistance

[4] Jacob (2004) and Chyn (2018) evaluate the impacts of public housing demolitions on outcomes of children using the same empirical strategy. Their analysis shows that displaced and non-displaced households have similar characteristics before building demolitions began.

[5] These welfare measures focus on adults and do not include any potential benefits of relocation for children in displaced households.

(e.g., Bergman, Chan and Kapor, 2020; Waldinger, 2021). In the work most closely related to our analysis, Galiani, Murphy and Pantano (2015) use data from the Moving to Opportunity (MTO) experiment to estimate a neighborhood choice model and study how mobility restrictions shape program participation and voucher take-up. To the best of our knowledge, their analysis is the only other prior study that also relies on data from a randomized or natural experiment to estimate a neighborhood choice model. While our empirical approach is broadly similar, we focus on estimating the welfare impacts of forced moves. Our analysis shows that rental vouchers generate substantively large welfare gains by expanding neighborhood choice. Notably, our analysis suggests that these benefits cannot offset moving costs and equilibrium adjustments in the housing environment that make displacement very costly for public housing residents.

3 Background

3.1 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.[6]

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 funding cuts during the 1980s that 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 Severely Distressed Public Housing, 1992).

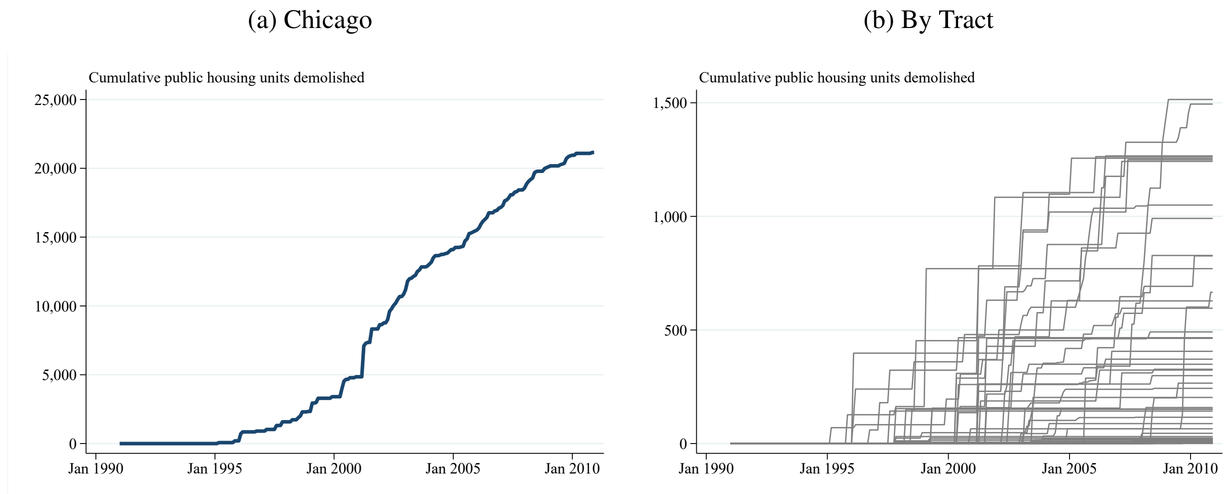
3.2 Chicago's HOPE VI Demolitions and Redevelopment

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).[7] Funding for demolition was provided through the

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

[7]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).

Figure 1: Time Series of Public Housing Demolitions in Chicago



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.

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.

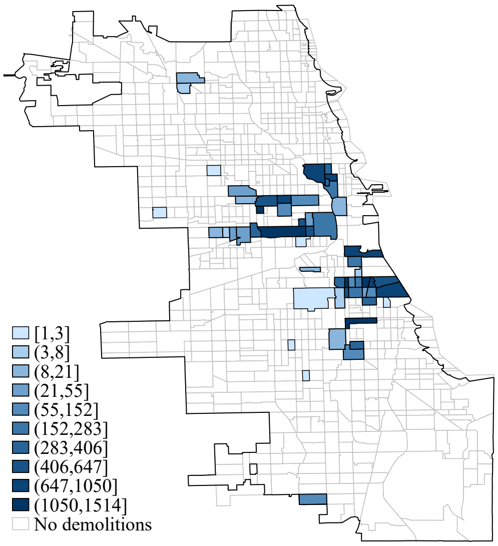
Chicago was one of the largest recipients of HOPE VI financing. From 1996 to 2003, the city received \$83.4 million in grant funding specifically for building demolitions (Aliprantis and Hartley, 2015). Public housing residents were evicted if their building was demolished and received offers for Section 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, 2023). 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.

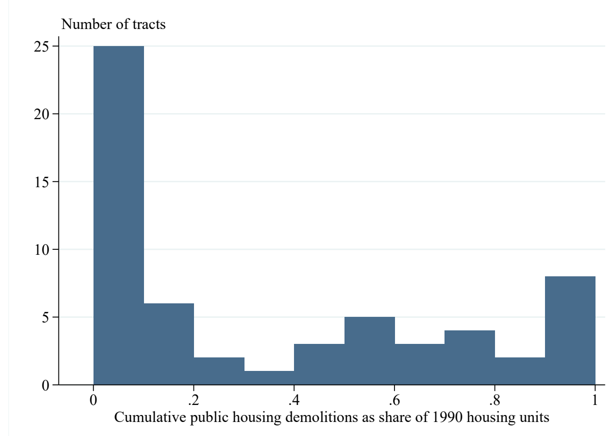
Figure 2 summarizes the spatial variation in public housing demolitions. In Panel A, we plot deciles of the cumulative number of demolitions in each tract between 1995 and 2010. There is considerable variation: 20 percent of tracts saw fewer than 8 public housing unit demolitions during this period, while 20 percent of tracts saw at least 647 units demolished. The consequences of these demolitions for neighborhoods likely depend 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 locations experienced demolitions that account for no more than 20 percent of the 1990 housing stock, but demolitions

Figure 2: Spatial Variation in Public Housing Demolitions in Chicago

(a) Cumulative Demolitions by Tract



(b) Distribution of Demolitions as Share of 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.

exceeded 50 percent of the 1990 housing stock in 37 percent of tracts with a 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, Kapur and Mithani, 2017).

Descriptive statistics from land-use data quantify the incomplete nature of redevelopment at former public housing sites.[8] 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).[9] Even by 2015, there was minimal additional progress as the share of vacant land stood at 35 percent.

[8]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.

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

4 Data

We compile data from two main sources for our analysis. First, we rely on Chicago Housing Authority (CHA) records from Sandler (2017) 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 a census tract 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.[10]

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 (Logan, Xu and Stults, 2014). 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 limited to tracts for which the key variables used in our analysis are not missing, contains 1,240 tracts in Cook County, Illinois.[11] On average, each tract has about 4,000 residents.

5 Motivating Facts: Public Housing Demolitions, Neighborhood Composition, and Housing Prices

This section documents descriptive facts on the relationship between HOPE VI demolitions and neighborhood outcomes. We show 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 key features of our spatial equilibrium model.

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 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.[12] 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

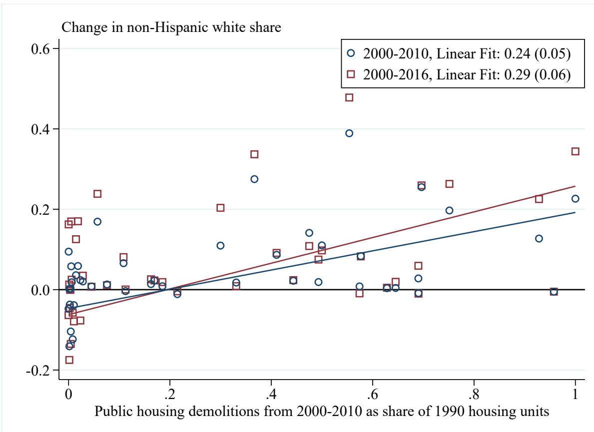
[10]We use 5-year ACS tabulations, covering 2008–2012 (which we refer to as 2010) and 2014–2018 (which we refer to as 2016). Population counts in 2010 come from the decennial census. Throughout, we use household counts and the race/ethnicity of the head of household.

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

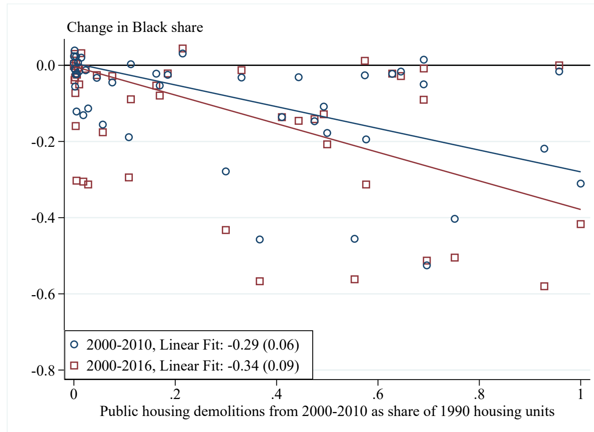
[12]Because all tracts with demolitions have a different amount of public housing demolitions as a share of 1990 housing units, this figure shows tract-level values for all tracts with demolitions and a separate average for all tracts with no demolitions.

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

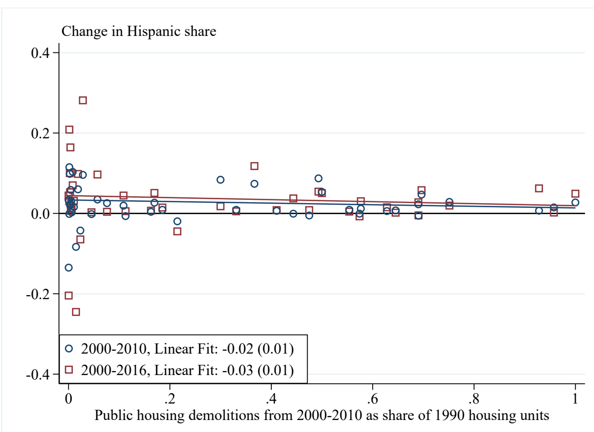
(a) Non-Hispanic White Population Share



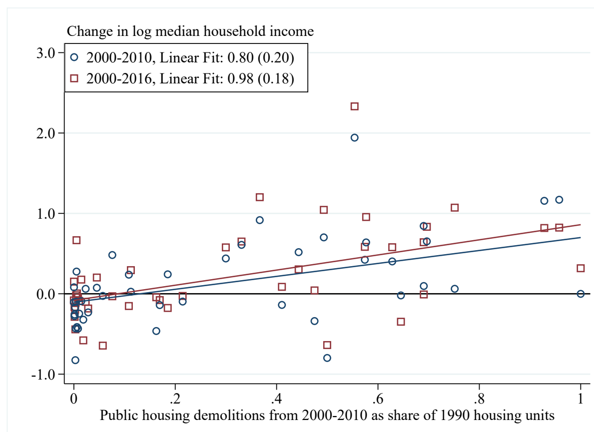
(b) Black Population Share



(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.

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 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, 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 demolitions.[13] 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 demolitions. 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.[14] 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.[15]

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 the different channels driving these neighborhood changes and quantify the associated consequences for households' welfare throughout Chicago. The basic ingredients of this model are motivated by the evidence presented above. Concretely, our model features heterogeneous preferences across demographic groups in a sorting framework, where both rents and demographics are equilibrium objects (Bayer and Timmins, 2005).

6 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 public housing or the socioeconomic and demographic characteristics of a neighborhood's residents. 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 (e.g., Bayer, Ferreira and McMillan, 2007; Davis, Gregory and Hartley, 2023) that features endogenous amenities with a model of housing supply.

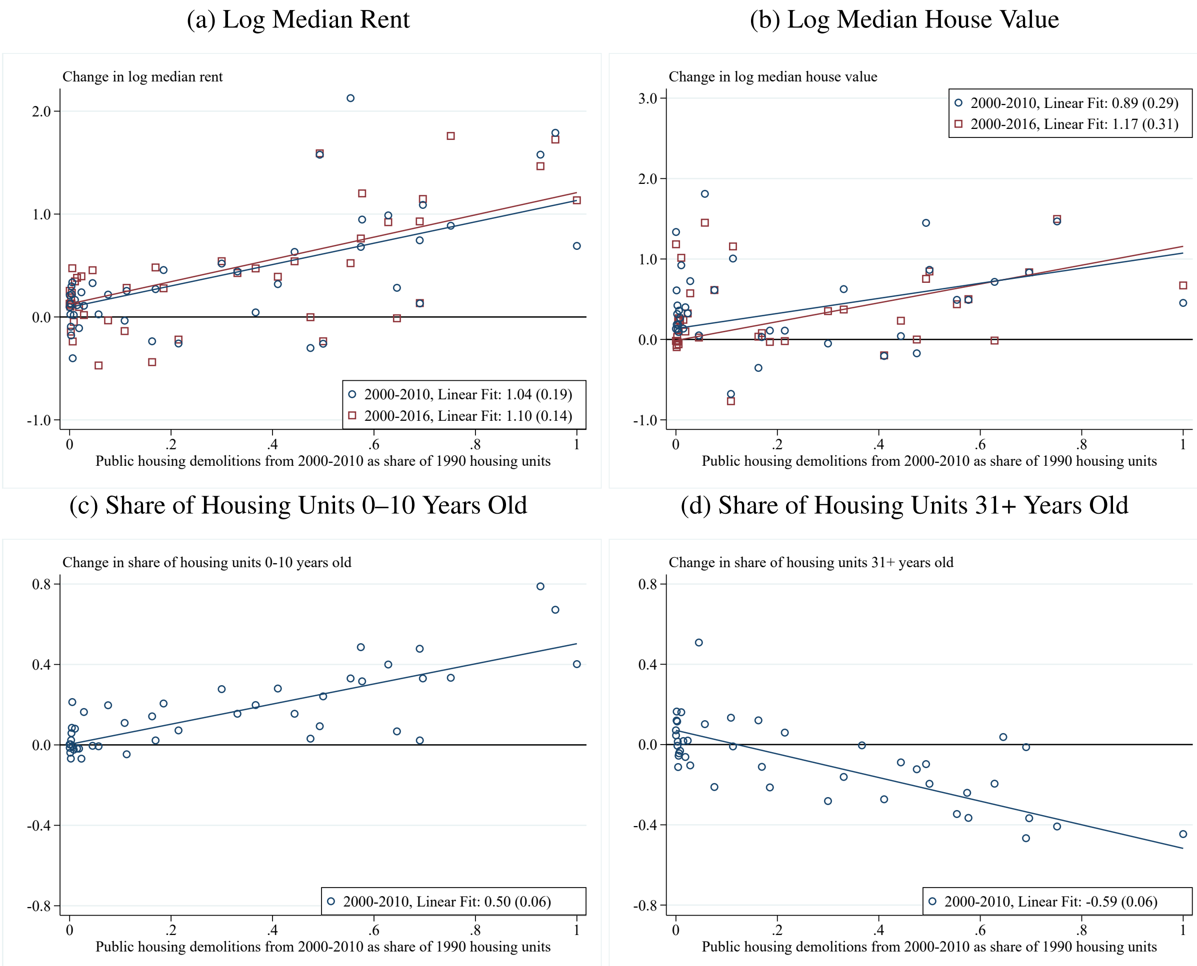
We begin by specifying a flexible model for household location choice as follows. Households of race-

[13] 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.

[14] 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.

[15] 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.

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



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.

by-income group k choose their neighborhood location j 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 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.

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 of households that want to live in j across groups $k = 1, \ldots, K$:

$$\mathcal{D}_{jt}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_{\mathbf{t}}; \alpha) = \sum_{k=1}^{K} N_{jt}^k = \sum_{k=1}^{K} \mathcal{P}_{jt}^k(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x}_t, \xi_{\mathbf{t}}^{\mathbf{k}}; \alpha^k) N_t^k, \tag{3}$$

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 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.

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 that hold for all neighborhoods $j = 1, \ldots, J$:

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

$$\frac{\mathcal{D}_{jt}^B(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_{\mathbf{t}}; \alpha) + N_{jt}^{PH}}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_{\mathbf{t}}; \alpha) + N_{jt}^{PH}} = b_{jt}^*, \tag{6}$$

$$\frac{\mathcal{D}_{jt}^H(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_{\mathbf{t}}; \alpha)}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_{\mathbf{t}}; \alpha) + N_{jt}^{PH}} = h_{jt}^*, \tag{7}$$

where $\mathcal{D}_{jt}^B(\cdot)$ and $\mathcal{D}_{jt}^H(\cdot)$ are the equilibrium number of Black and Hispanic households in neighborhood j, and N_{jt}^{PH} is the number of public housing residents in neighborhood.[16] 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

[16] In our main analysis of the city-wide effects of demolitions, we assume that the location choice of households in public housing is fixed and set each neighborhood's public housing population to the level observed in time period t. In Section 10, we estimate a location choice model for public housing residents and use these results to provide welfare estimates for this population.

for understanding the effects of demolitions on neighborhoods.[17]

7 Quantification of the Model

7.1 Demand Estimation

To study the consequences of public housing demolitions using our model, a necessary step is to obtain estimates of the household preference parameters. Since the choice probabilities determined by equation (2) are invariant to scale, we normalize the indirect utility of the outside option, which is assumed to be the choice of living outside of Cook County within Illinois and indexed as $j = 0$, to be equal to zero: $\delta_{0t}^k = 0$ (Train, 2009).[18] This normalization also ensures that our identifying variation comes from changes in neighborhoods *within* Cook County.[19] 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 explicitly focus on this definition for poor households given that they were more likely to live in public housing neighborhoods and thus were more exposed to changes triggered by demolitions.[20,21]

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.[22] Second, due to the finite sample nature of the ACS data, 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

[17] Appendix B describes the numerical procedure that allows us to solve for the equilibrium with endogenous amenities.

[18] Due to this scale-invariant property, under the previous normalization, δ_{jt}^k only identifies utility differences with respect to the outside option.

[19] The inclusion of this outside option makes our model 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. Counterfactual simulations are identified as long as the utility level outside Cook County does not change in response to demolitions inside Chicago (Kalouptsidi, Scott and Souza-Rodrigues, 2021).

[20] 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.

[21] When choosing the number of income bins we face a trade-off between adding more heterogeneity and finite-sample bias due to sparsity (Dingel and Tintelnot, 2020).

[22] 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).

weighted average of our frequency estimates across census tracts:

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

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:

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

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.[23] 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.

7.2 Identification of Preference Parameters

The system of equations (5)–(7) shows that when estimating equation (10) both rents and demographic composition are a function of unobservable demand shocks, ξ^k_{jt}. For example, unobservable housing quality leads to higher house prices, and ignoring this confounder would upward bias the coefficient for prices. We deal with this identification threat in two ways.

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^k_j, 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 inside Cook County relative to the outside option by including year t fixed effects, λ^k_t.[24] Both fixed effects vary arbitrarily by race-and-income group. After the inclusion of these fixed effects, our empirical demand model is given by:

$$\begin{aligned} \log\left(\frac{\tilde{P}^k_{jt}}{\tilde{P}^k_{0t}}\right) &= \alpha^k_p \ln(p_{jt}) + \alpha^k_b b_{jt} + \alpha^k_h h_{jt} + \alpha^k_{Inc} \ln(Inc_{jt}) \\ &\quad + \alpha^k_{PH} PH_{jt} + \theta^k x_{jt} + \lambda^k_j + \lambda^k_t + \tilde{\xi}^k_{jt}. \end{aligned} \tag{10}$$

[23] 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.

[24] Note that shocks common across all tracts in Illinois, such as changes in inflation rates, are differenced out in equation (10).

Second, we extend the instrumental variable approach proposed in Bayer, Ferreira and McMillan (2007) and Berry, Levinsohn and Pakes (1995) to a setting with repeated cross-sections. To implement our approach, we begin by constructing instruments using measures of public housing shares and housing characteristics in neighborhoods that are farther than 3 miles away. In our 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 shown by equation 2. As substitutes, shifts in the characteristics of other neighborhoods alter equilibrium prices (Berry, Levinsohn and Pakes, 1995). 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. In this way, we argue that the approach proposed by Bayer, Ferreira and McMillan (2007) also creates exogenous variation that identifies preferences over demographics, as long as different demographic groups also differ in their willingness to pay for housing characteristics.

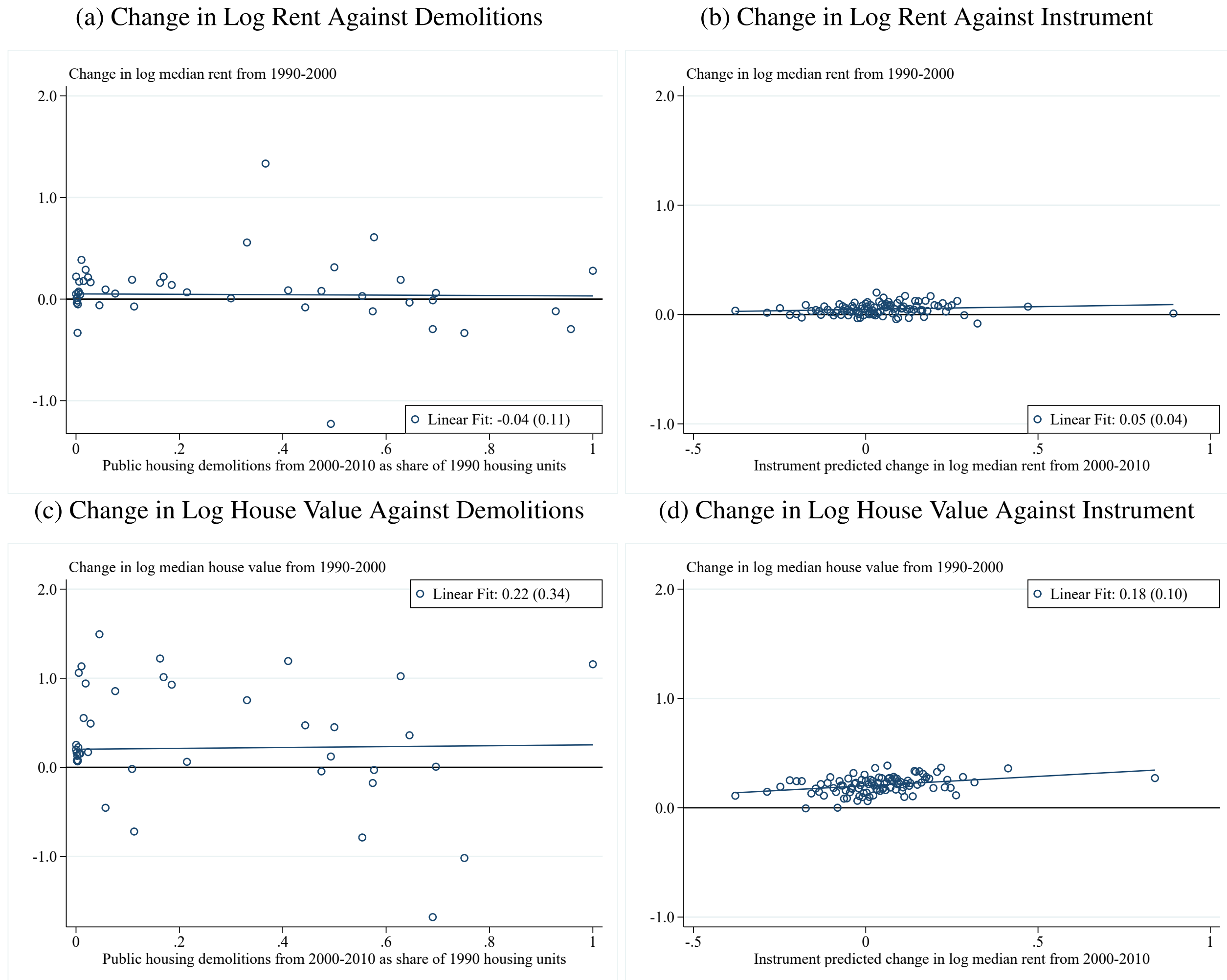
More formally, given the repeated cross-section nature of our data, our exclusion restriction only requires that *changes* in the public housing share and physical housing characteristics in distant neighborhoods are uncorrelated with *changes* in unobservable neighborhood characteristics, conditional on changes in observable characteristics:

$$\mathbb{E}[\Delta\tilde{\xi}^k_{jt}\Delta z_{jt}|\Delta PH_{jt}, \Delta x_{jt}, \Delta\lambda^k_t] = 0,$$

which is an identifying assumption similar to Couture and Handbury (2020). Given our use of distant neighborhoods to construct instruments, our exclusion restriction is satisfied if changes in exogenous characteristics of neighborhoods that are at least 3 miles away are uncorrelated with unobservable time-varying demand shocks $\Delta\tilde{\xi}^k_{jt}$ conditional on changes in observables.

In the spirit of the optimal instrument literature (e.g., Chamberlain, 1987; Reynaert and Verboven, 2014), we use a three-step approach following previous studies (Bayer, Ferreira and McMillan, 2007; Wong, 2013; Calder-Wang, 2019; Anagol, Ferreira and Rexer, 2021; Allen, Arkolakis and Takahashi, 2020). First, we estimate preference parameters using equation (10) and the already-discussed instruments. Second, we solve for equilibrium rents (p_{jt}) and location choices ($\mathcal{P}^k_{jt}$) under the assumption that there are no unobserved time-varying determinants of neighborhood quality (i.e., $\tilde{\xi}^k_{jt} = 0$). Using the location choices, we construct the share of each neighborhood that is composed of each race-by-income group k. Finally, we estimate equation (10) using as instruments the simulated price and the share of households in each race-by-income group. In summary, this three-step approach increases the power of our instrumental variable approach in two ways. First, it includes information from the equilibrium market clearing conditions from equations (5)–(7). Second, it concentrates variation from many instruments into functions that are especially relevant for the endogenous variables of interest.

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



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.

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 related to poor building maintenance which led housing authorities to select particular buildings for demolition. Second, changes in the housing stock in distant tracts (e.g., due to redevelopment) were unlikely to depend on unobservable trends in a given tract. 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).

7.3 Household Preference Results

Table 1 presents instrumental variable estimates of equation (10).[25] 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). 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 failing 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.[26]

There are two main caveats for the interpretation of the results in Table 1. First, we interpret these estimates as reduced-form parameters that may reflect the combined impact of additional preferences that we do not explicitly model (Caetano and Maheshri, 2021; Davis, Gregory and Hartley, 2023). 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 or lower crime levels (Almagro and Domínguez-Iino, 2024; Khanna et al., 2023). 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. Notably, our parametric framework allows for several equilibrium responses through rents and demographic composition while keeping tractability that only requires aggregate census data for its quantification. 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, 2022).

To compare preferences across groups, we calculate the implied willingness to pay for different neighborhood characteristics.[27] Poor White residents are willing to increase their annual rent by about \$44 and \$36 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 \$134 and \$135 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-

[25] 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.

[26] 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.

[27] Coefficients in Table 1 are measured in utils and thus cannot be compared across groups.

by-income groups have positive willingness to pay for reductions in the public housing share.

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, Ferreira and McMillan (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.

How do these estimates compare to previous studies? One comparison for our analysis comes from Galiani, Murphy and Pantano (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 characteristics (log median number of rooms, log median 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).[28] 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. 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.

[28] 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.

7.4 Housing Supply

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 (2024), 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 the implied equilibrium quantities under the estimated demand preferences, $\hat{\alpha}$, when we remove unobserved demand shocks, $\tilde{\xi}_{jt} = 0$:[29]

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

7.5 Assessing Model Fit

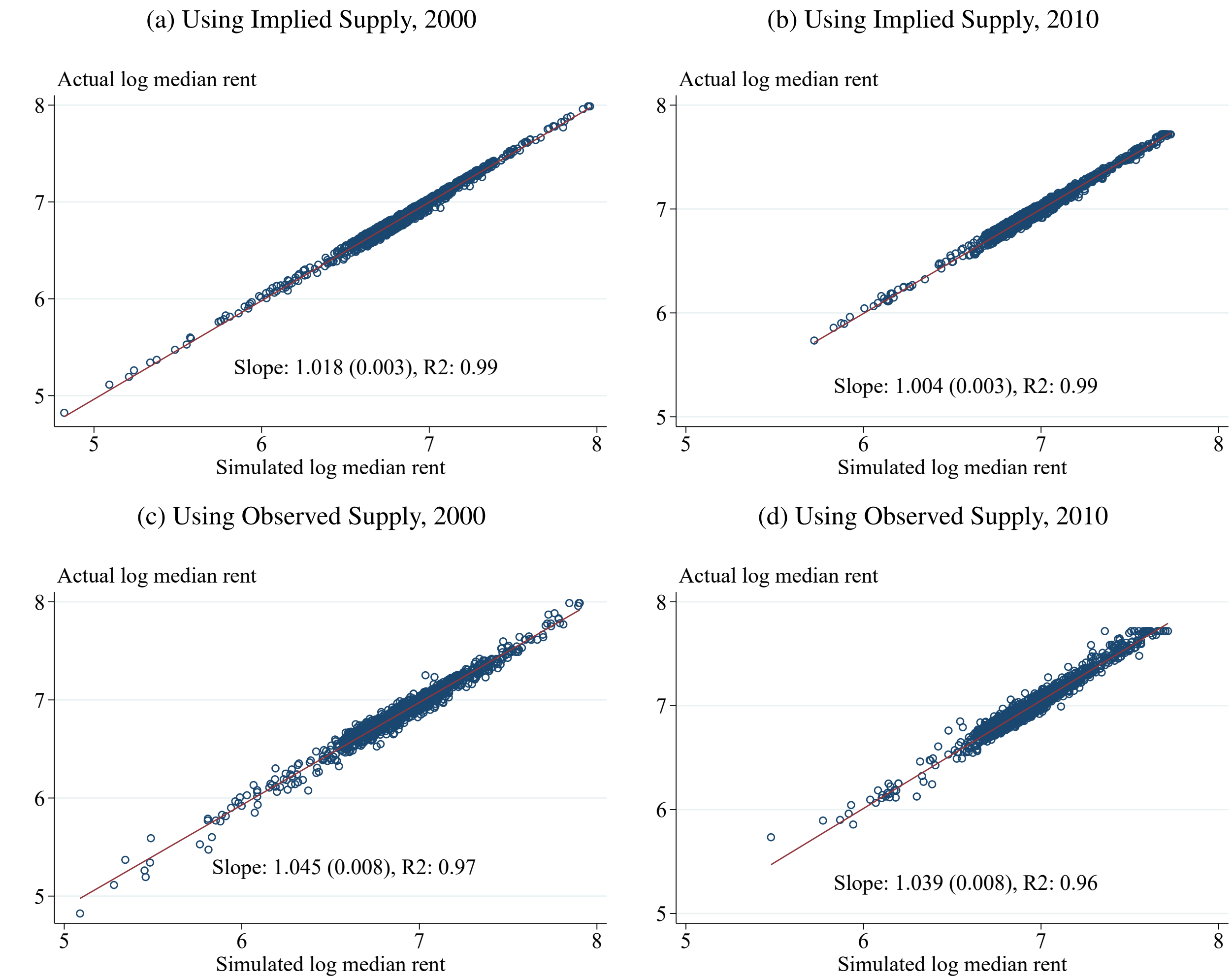
Before describing the welfare consequences of public housing demolitions, we conduct one in-sample and one out-of-sample validation exercise to assess how well our model fits equilibrium rental prices. In both exercises, we focus on the explanatory power of the explicitly-modeled elements by setting the unobserved time-varying index of neighborhood quality equal to zero (i.e., $\tilde{\xi}_{jt}^{k} = 0$ for all j, t, and k). Rents 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 following Section 7.4. As a result, actual and simulated rents in these panels differ only because of 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 in 2010. These results demonstrate a high degree of in-sample 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. We also assume that the housing supply shifter, $\hat{\theta}_{jt}$, 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

[29]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}. Our approach has the advantage that no unobservable demand components enter the calibrated supply shifters, $\hat{\theta}_{jt}$.

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



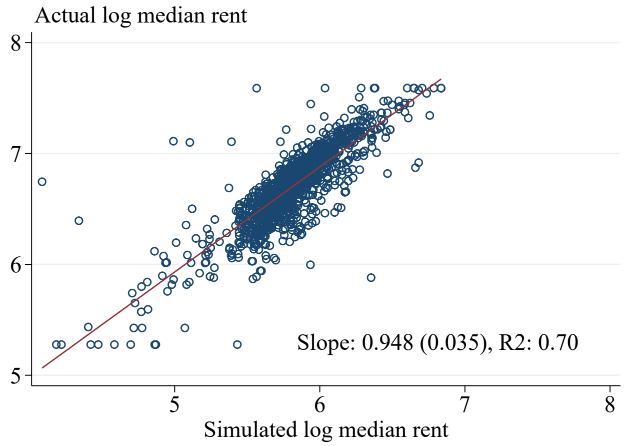
Notes: This figure plots actual log rents in census tracts against log rents that are implied by the model estimates where unobservable components of neighborhood quality are set to zero (i.e., $\tilde{\xi}_{jt}^{k} = 0$ for all k, j, and t). In Panels A and B, the number of housing units supplied is 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.

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.

8 Impacts of Public Housing Demolitions

What impact did public housing demolition have on welfare? To answer this, we use our estimated demand parameters and our calibrated supply curve to compute welfare in several counterfactual scenarios using the model and equilibrium conditions from Section 6. After describing our main results, we consider alternative models in Section 8.3. Specifically, Section 8.3.1 introduces moving costs, Section 8.3.2 analyzes the possibility of multiple equilibria, Section 8.3.3 allows for intensive margin adjustments in housing

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



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.

consumption, and Section 8.3.4 evaluates the effect of displaced public housing residents on market rents.

8.1 Measuring Welfare

Throughout the text, we focus on household welfare for residents of Cook County unless otherwise specified. Using the estimated preference parameters, 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:

$$\mathcal{CS}^k(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi^{\mathbf{k}}; \alpha^k) = \log\left(\sum_{j=1}^{J} \exp\left(v_{jt}^k(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi^{\mathbf{k}}; \alpha^k)\right)\right) + \Gamma^k,$$

where $v_j^k(\cdot)$ is indirect utility specified in equation (1). We assume that the unobservable demand factor is equal to its conditional mean across all groups, (i.e., $\tilde{\xi}_{jt}^k = 0$ for all j, t, and k), and we suppress that term inside our consumer surplus measure to simplify notation.[30] 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.

To compute renter welfare changes from scenario $(\mathbf{p}^1, \mathbf{b}^1, \mathbf{h}^1, \mathbf{x}^1)$ relative to scenario $(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^0)$ in monetary terms, we define the group-specific rent equivalent, RE^k, following Small and Rosen (1981):

$$RE^k \equiv \frac{\mathcal{CS}^k(\mathbf{p}^1, \mathbf{b}^1, \mathbf{h}^1, \mathbf{x}^1; \alpha^k) - \mathcal{CS}^k(\mathbf{p}^0, \mathbf{b}^0, \mathbf{h}^0, \mathbf{x}^0; \alpha^k)}{\lambda^k},$$

[30] Alternatively, we could incorporate the estimate $\hat{\tilde{\xi}}_{j,t}^k$ for $t = 2000, 2010$ into the utility function. This would require the realization of this component to remain unchanged across different scenarios, which is arguably a stronger assumption. It is worth noting that results are qualitatively and quantitatively similar when we do so.

where λ^k is the marginal utility of housing consumption for group k.[31] Note that positive values of the rent equivalent are associated with higher welfare in the counterfactual world.

Given that welfare changes in rent equivalent units are measured in dollars, we can readily measure the welfare change for homeowners as the sum of the rent equivalent and the rental income that accrues to them. 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:

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

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

8.2 Main Results

We simulate our counterfactuals fixing $t = 2010$. Our goal is to compare welfare from counterfactual scenarios to the baseline observed scenario where the CHA destroyed public housing units.[32] 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 2010 in which the public housing share in each tract is held constant at its level in 2000.[33] We solve for the equilibrium rents, demographic shares, and households' location choices in this counterfactual 2010.

Figure 8 reports the welfare changes in rent-equivalent units for the average household due to public housing demolitions for each group.[34] 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 \$285 increase in their annual rents (representing a 2 percent increase relative to the mean annual rent in Cook County in 2010). Poor White households experience essentially no change in utility due to public housing demolitions. In contrast, Black and Hispanic households are worse off because of demolitions. The decrease in utility is equivalent to -\$247 per year for poor Black households and -\$429 for non-poor Black households, with Hispanic households experiencing a decrease in well-being that is smaller but still sizable. 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

[31] We define marginal utility of housing consumption for group k as follows:

$$\lambda^k \equiv -\frac{\alpha^k_p}{\sum_j \frac{N^k_j}{N^k} p_j}.$$

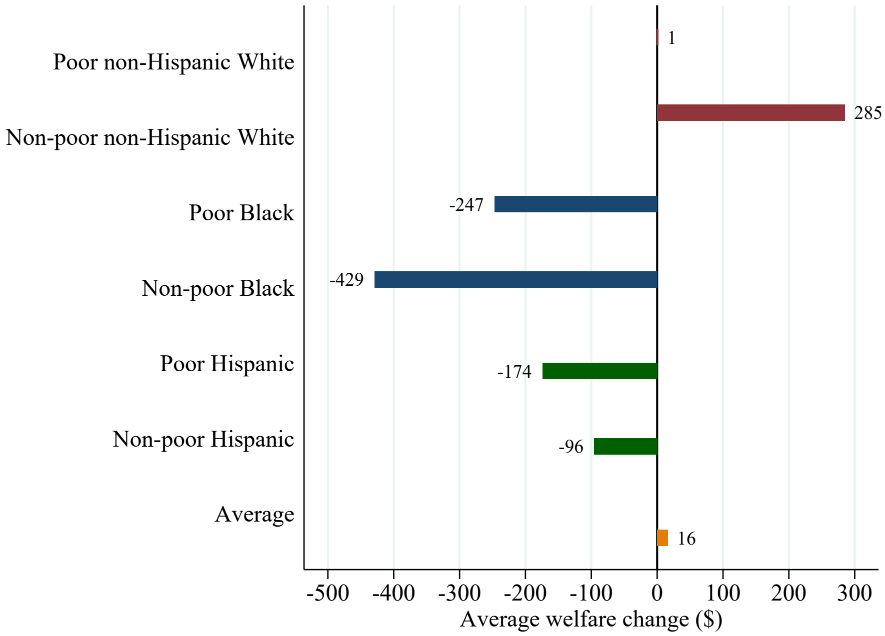
The marginal utility of consumption for Non-Hispanic White Poor households is 3.3e-5, which means that a dollar translates into 3.3e-5 utils. For the rest of the groups λ^k is equal to 4.0e-6, 1.8e-5, 2.7e-5, 1.8e-5 and 7.4e-6 for Non-Hispanic White Non-Poor, Black Poor, Black Non-Poor, Hispanic Poor, and Hispanic Non-Poor households respectively.

[32] Because we compare counterfactual 2010 outcomes to actual 2010 outcomes, these results are not mechanically influenced by changes from 2000 to 2010 in Cook County relative to the outside option.

[33] In the counterfactual scenario without public housing demolitions, we remove the number of poor Black households in 2010 by the number of occupied public housing units that were demolished (approximately 15,000) from the market and assign them to the counterfactually non-demolished public housing units.

[34] We also estimate the welfare effects for other race/ethnicity households. The change in rent equivalent welfare is -\$643 and -\$255 for poor and non-poor other race/ethnicity households, respectively.

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.

average welfare of non-Hispanic White, Black, and Hispanic households by a negligible amount ($16) due to the fact that non-poor White households are the largest demographic group in our context.

8.2.1 Decomposition Analysis

To understand the overall impact of public housing demolitions on welfare, Table 2 provides decomposition results. This exercise selectively highlights the quantitative importance of various equilibrium channels that are embedded in our model. We start in the counterfactual 2010 scenario without demolitions and turn on each margin of adjustment step-by-step, as indicated by each row label of Table 2.[35] Intermediate scenarios only allow some endogenous features to adjust, and thus are partial equilibrium results.[36]

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 counterfactual 2010 levels with public housing. Because all groups view public housing as a disamenity, the rent

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

$$\frac{1}{\lambda^k}\Big(\mathcal{CS}^k(\mathbf{p}^0,\mathbf{b}^0,\mathbf{h}^0,\mathbf{x}^{demolitions};\alpha^k)-\mathcal{CS}^k(\mathbf{p}^0,\mathbf{b}^0,\mathbf{h}^0,\mathbf{x}^0;\alpha^k)\Big),$$

where $\mathbf{x}^{demolitions}$ differs from $\mathbf{x}^0$ only by incorporating public housing demolitions in each tract.

[36] The "all channels" results in Panel A row 4, as well as the results in Panel B and C, are based on a general equilibrium analysis.

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.[37] 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 re-sorting 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 re-sorting 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 this scenario as our simulations find that demolitions increased rents substantially. Overall, these results show how demolitions led to unequal impacts on renters due to changes in demographic composition and the cost of housing.

We summarize the effects on homeowners in Panel B. As noted in Section 8.1, renters and homeowners have the same preferences, but the welfare of the latter is also 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 compared to renters.

Finally, the remaining rows of Table 2 summarize both the average 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.[38] Due to this disparity in homeownership, the rent equivalent welfare gain of non-poor White households is much higher ($285 versus $30) and the gap between this group and poor Black households increases by 58 percent.

[37] 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).

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

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

	Change from baseline (2010 Census)					
	Non-Hispanic White		Black		Hispanic	
	Poor	Non-poor	Poor	Non-poor	Poor	Non-poor
Counterfactual scenario	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Results for renters						
Destroy buildings in tract	144	165	192	392	193	199
... and change neighborhood composition via demolitions	173	340	82	-145	96	61
... and change neighborhood composition via resorting	178	363	56	-244	80	44
... and change housing prices (all channels for renters)	-160	30	-306	-598	-266	-294
Panel B. Results for homeowners						
... and redistribute rents to homeowners (all channels for owners)	154	344	8	-284	48	20
Panel C. Full equilibrium results						
Average welfare change across renters & owners	1	285	-247	-429	-174	-96
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.

8.2.2 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 5. 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.

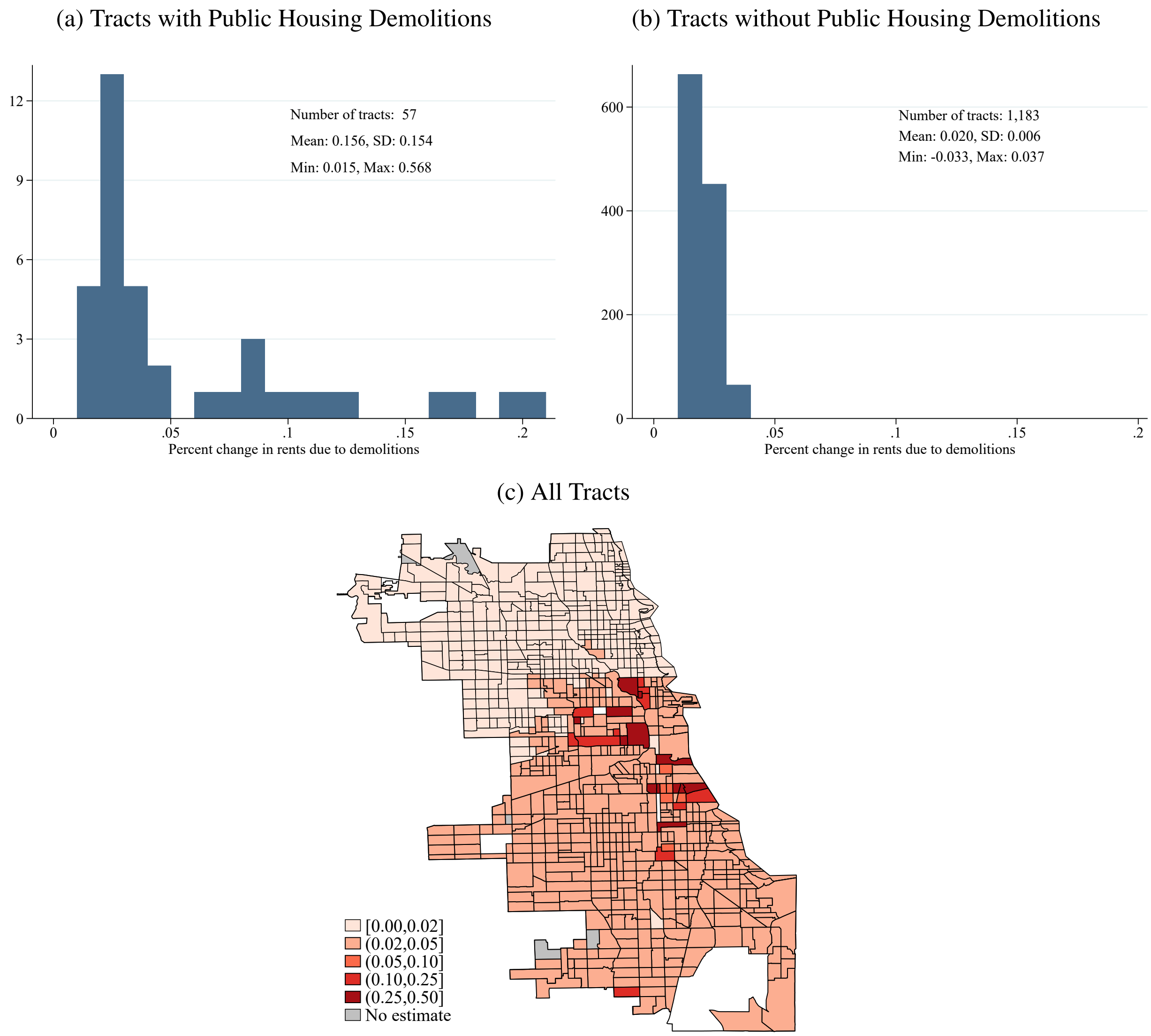
Theoretically, our model suggests that there will be heterogeneous effects of public housing demolitions across neighborhoods. Given the household preference estimates, we expect rents to go up in neighborhoods which had demolitions. For neighborhoods without demolitions, there are two potentially offsetting effects that arise through substitution across neighborhoods and equilibrium forces. First, demand might shift toward neighborhoods that directly 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 demolitions. Tracts with a demolition saw an average rent increase of 15.6 percent (Panel A). The distribution of changes for these tracts exhibits a fat right tail. The areas with the largest rent changes had the most extensive number of demolitions. The 15.6 percent effect for tracts with demolitions is almost 7 times larger than the average rent increase of 2.0 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.[39] 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

[39]In Appendix Figure A.2, we plot the tract-level percent change in 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.

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

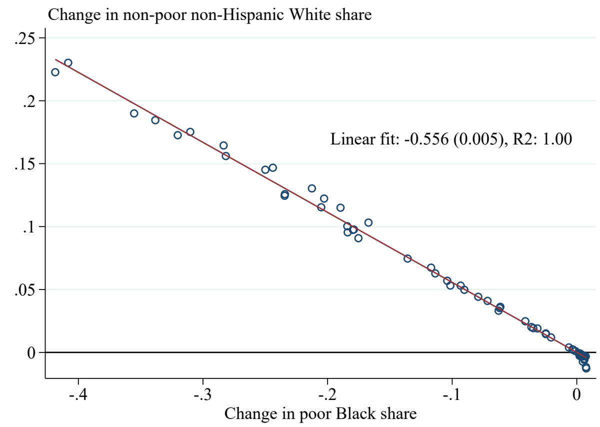


Notes: This figure displays the distribution of the percent change in median rents due to public housing demolitions. We construct this change using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario 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 3 tracts with a change below 0. 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 percent change in median rents.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

that are within 0.1 miles of a demolition site, but only 1 percent for neighborhoods that are 20 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 for each tract in Chicago. The slope coefficient is precisely estimated and implies that areas where demolitions caused a 1 percent decrease

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 U.S. Census Bureau.

in the poor Black household share experienced a 0.56 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 Appendix Table A.8, our estimates imply that 72 percent of the aggregate increase in rents comes from neighborhoods without public housing demolitions.[40]

8.3 Robustness

8.3.1 Incorporating Moving Costs

One potential concern is that our main analysis does not account for bilateral moving costs. A number of prior studies in the literature have considered the role that moving costs play in location choice and

[40] Put differently, the 5 percent of neighborhoods with a public housing demolition account for 28 percent of the city-wide housing price increase.

sorting models (Kennan and Walker, 2011; Bayer, Keohane and Timmins, 2009; Ferreira, 2010). To quantify moving frictions, these studies frequently rely on longitudinal data that records an individual's location over time to measure migration flows. Aggregated at the location level, these studies use information on the likelihood of moving to a new location to quantify mobility frictions, while still relying on within-neighborhood variation to estimate preferences over location characteristics. Although we cannot track how individual households move across decades due to the cross-sectional nature of the Census data available for our analysis and cannot directly estimate moving costs, we can show how our welfare conclusions can potentially change in the presence of an assumed moving costs.

To shed light on how moving costs potentially affect our welfare conclusions, we extend our model to incorporate mobility frictions into the framework from Section 6. Because data limitations prevent us from directly estimating moving costs, we instead calibrate the utility loss from having to move to a different location. Specifically, we rely on an estimated moving cost of \$38,000 from Bayer et al. (2016) which is based on a dynamic neighborhood sorting model.[41] We incorporate these moving costs into an extended version of our model that follows the approach from Galiani, Murphy and Pantano (2015) which estimates a static (within city) neighborhood choice model featuring mobility costs.

Formally, we assume that a household i of type k is endowed with an original location, denoted by j^k_{it-1}, and decides whether to move to a new neighborhood at the start of period t. Allowing for a mobility cost MC, we now assume a utility flow of relocating to location j is given by:

$$\delta^k_{jt} + \epsilon^k_{ijt} - MC \cdot \mathbb{1}\{j \neq j^k_{it-1}\},$$

where δ^k_{jt} is the component of indirect utility for neighborhood j that is common to all households of group k and ϵ^i_{jt} is a type I Extreme Value idiosyncratic shock. Given this assumption, the probability of moving to j conditional on origin j^k_{it-1} is given by:

$$\mathcal{P}^k_{jt|j^k_{it-1}}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi^{\mathbf{k}}_{\mathbf{t}}; \alpha^k) = \frac{\exp\left(\delta^k_{jt} - MC \cdot \mathbb{1}\{j \neq j^k_{it-1}\}\right)}{\sum_{j'} \exp\left(\delta^k_{j't} - MC \cdot \mathbb{1}\{j' \neq j^k_{it-1}\}\right)}.$$

To use this framework in our HOPE VI context, we define period $t-1$ to be the year 2000 (the period before the public housing demolitions we study) and set the period t equal to the year 2010. Given that Bayer et al. (2016) estimate a yearly moving cost, we calibrate a ten-year moving cost by discounting their estimates by a ten-year discount factor.[42] Our approach also requires that we specify the original distribution of locations for individuals, which we set to match our simulated pre-demolition distribution for the year 2000:

$$\breve{\mathcal{P}}^k_{j,2000} \equiv \mathcal{P}^k_{j,2000}(\mathbf{p}_{2000}, \mathbf{b}_{2000}, \mathbf{h}_{2000}, \mathbf{x}_{2000}, \xi^k_{2000}; \alpha^k).$$

In this extension of the model, our demand function is now computed as follows:

$$\mathcal{D}^{MC}_{jt}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_t; \alpha) = \sum_k N^k_t \sum_{\ell=0}^{J} \mathcal{P}^k_{jt|\ell}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi^{\mathbf{k}}_{\mathbf{t}}; \alpha^k)\breve{\mathcal{P}}^k_{j,2000}(\ell), \tag{11}$$

[41] We infer an estimated moving cost using the results from Table 4 in Bayer et al. (2016) in combination with summary statistics for Chicago on the average housing value in 2000 dollars (which we estimate to be \$206,000).

[42] Because they use a one-year discount factor of 0.95, the compounded annual discount factor after ten years is 0.95^{10}.

where $\breve{\mathcal{P}}^k_{j,2000}(\ell)$ denotes the ℓ^{th} component of vector $\breve{\mathcal{P}}^k_{j,2000}$.The equilibrium conditions are defined by a similar system of equations as specified in Section 6 using the demand estimates that account for mobility costs MC.

Theoretically, the introduction of moving costs makes households less responsive to changes in neighborhood characteristics. This reduced responsiveness can change equilibrium rents in two ways. On the one hand, households have lower incentives to move away from their original neighborhood, which means that they are willing to pay a higher rent to live there. On the other hand, households in *other* neighborhoods also have less incentives to move to a different neighborhood, which lowers the rents in these areas. These two effects go in opposite directions and it is hard to know a priori which effect dominates. In turn, this makes it difficult to even sign the potential biases in welfare results in the absence of moving costs.

Overall, the results based on our augmented model show that our main finding of racial disparities in the welfare impacts of public housing demolitions are largely qualitatively and quantitatively robust. Appendix Table A.9 reports results from the baseline and mobility cost versions of our model in Panels A and B, respectively. The final row of Panel B shows that the impacts of demolition remain similar when accounting for moving costs, with the largest increase in welfare for non-poor White households and the worst outcomes for Black households. The effects of demolitions tend to be more beneficial for White households in the presence of moving costs and more harmful for Black households.

What drives this pattern of results? A key input into the welfare effects is the change in house prices. When incorporating moving costs, we find that rents across the city increase by 3.3 percent after public housing demolitions. This is larger than the price increase in the absence of moving costs (2.5 percent), which is driven by households' decreased willingness to relocate to other neighborhoods in the presence of moving costs which in turn makes them more inelastic.[43] The greater price appreciation tends to make renters worse off and homeowners better off, which interacts with baseline differences in homeownership rates in producing disparate outcomes. Panel A of Appendix Figure A.4 shows that these results are very robust to using one-year moving costs that range between 0 (our baseline model) and double the $38,000 estimate from Bayer et al. (2016). Panel B shows that, in contrast, the increase in rents in neighborhoods with a demolition is sharply curtailed once one-year moving costs increase beyond about $50,000; nonetheless, the total effect on citywide rents is quite robust, which reflects the importance of neighborhoods without demolitions in generating aggregate results.

8.3.2 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

[43] The larger city-wide increase in rents is driven by neighborhoods without a public housing demolition. In these neighborhoods, rents increase by 2.9 percent after demolitions (compared to 2.0 percent without moving costs). In neighborhoods with demolitions, rents increase by 11.6 percent (compared to 15.6 percent without moving costs).

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 groups.

8.3.3 Allowing for Housing Consumption

Our baseline approach assumes that each household consumes one unit of housing and does not explicitly model how households choose how much quantity of housing they want to consume within a location. To explore the importance of this channel, we instead assume that households consume their optimal amount of housing under Cobb-Douglas preferences in addition to selecting the neighborhood in which they reside (e.g., Baum-Snow and Hartley, 2020; Davis et al., 2021).[44] Under these assumptions, the coefficient in the indirect utility function on the log housing price variable is equal to the housing expenditure share.

We construct housing expenditure shares separately for each race and income group using 2000 Census microdata and then calibrate the coefficient on price to be equal to such expenditure share. In practice, we difference the price component of indirect utility from the log choice probability. For this exercise, we use the log of median rents divided by the median number of bedrooms in each tract as our measure of housing prices, where the housing consumption decision is made on the number of rooms.[45]

In terms of the estimation, the Cobb-Douglas-based approach is not subject to the same concern over the endogeneity of rents since the model implies that the coefficients on rents should be calibrated in a prior step. As a result, this approach imposes less structure on the identification assumptions and the validity of the instruments.[46] In other words, the choice between our main approach and the Cobb-Douglas model implies a tradeoff between making identifying assumptions in our instrumental variable strategy and imposing assumptions about housing consumption.

In Appendix Table A.10, we compare the implied willingness to pay for neighborhood demographic characteristics and the presence of public housing from our baseline model (Panel A) and the Cobb-Douglas model (Panel B). The estimation results from the Cobb-Douglas model are broadly similar to those from our preferred model that abstracts from within-neighborhood housing consumption changes. Many of the

[44] Thus, between our baseline model in which households cannot adjust the amount of housing consumption, and the Cobb-Douglas model in which households can adjust housing consumption quite flexibly, we cover a range of potential models of housing consumption.

[45] Unfortunately, the decennial census data do not provide information about square footage.

[46] We still instrument demographic composition using the approach outlined in Section 7.2.

estimated willingness to pay parameters are also quite close in quantitative terms. Overall, these results suggest that our abstraction from intensive-margin housing consumption changes does not unduly influence the estimation results.

While the results from our willingness-to-pay analysis are reassuring for our main approach, we note two drawbacks associated with a Cobb-Douglas utility over housing consumption. First, we find relatively weaker evidence in favor of this approach in our out-of-sample validation exercises for 1990 data using the same procedure used for Figure 7.[47] As shown in Appendix Figure A.5, the out-of-sample fit of the Cobb-Douglas model is meaningfully worse than of the benchmark model (the R^2 falls from 0.70 to 0.42). We take this as evidence that our baseline model is relatively better able to capture the economic mechanisms of how people are consuming and sorting within the city. Second, another reservation is that the Cobb Douglas framework relies on more model assumptions about the demand structure when it comes to the estimation of preference parameters. By contrast, our main estimation implicitly treats housing consumption in a reduced form way, as we allow each coefficient on rent to vary by group with the remaining margins of within-neighborhood housing consumption being part of the group-specific location residual. The theoretical restrictions imposed by assuming Cobb Douglas preferences may be a potential explanation for its weaker performance in our validation exercises.

8.3.4 Assessing Pecuniary Externalities from Public Housing Households

In our main counterfactual exercise, we assumed that the households inside public housing were not part of the market. However, after demolitions these households could potentially create extra pressure on rents from an increase in demand, that could lead to "pecuniary externalities" of public housing demolitions. To asses the importance of these effects, we also simulate a scenario where we leave unchanged the number of poor Black-households in the market before demolitions. Because the number of residents displaced from public housing is small relative to the total number of residents in Chicago, namely less than 0.6%, results are very similar when the number of poor-Black households in the rental market remains unchanged before and after demolitions. Concretely, 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 exceeds 0.99, with a maximum absolute deviation of 3.6 percent. These small changes suggest that pecuniary externalities from public housing demolitions were limited relative to other effects in the case of Chicago.

9 Evaluating 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 improving 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—an exercise motivated by prior work highlighting the importance of building constraints

[47] Note that we assume that supply here is given in the total number of rooms offered in a neighborhoods as opposed to the total number of housing units.

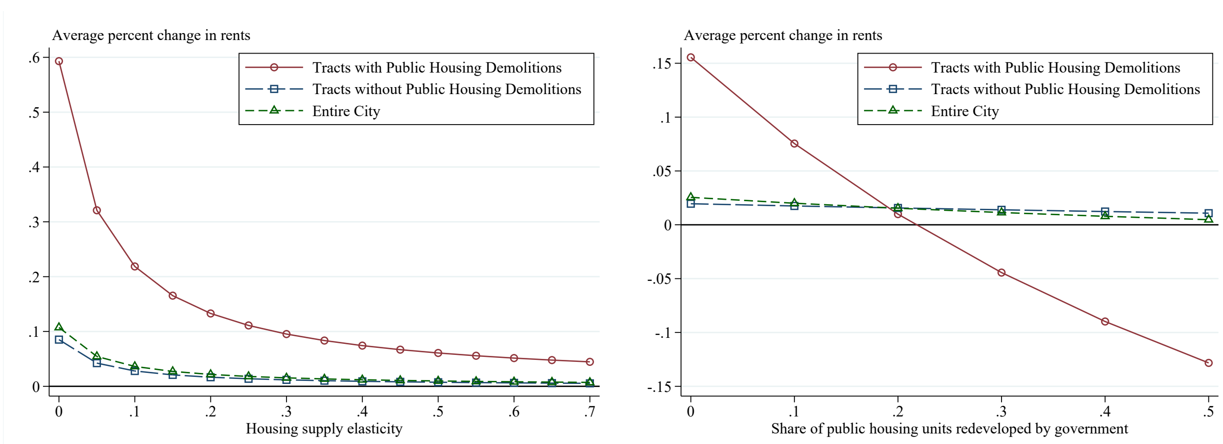
(Gyourko, Saiz and Summers, 2008; Saiz, 2010). Second, we study how the scale of redevelopment in public housing areas matters by estimating counterfactuals in which we assume additional market rate units are created in neighborhoods which featured demolitions. This second policy is motivated by the fact that the original HOPE VI plan also included the redevelopment of affordable housing units that would have replaced the demolished public housing projects.

9.1 Welfare Effects under Alternative Supply Elasticities

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 (2024). 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 supply elasticity would reduce the positive price impacts that particularly harm poor households.

Figure 11: Consequences of Public Housing Demolitions on Housing Prices Under Alternative Housing Policies

(a) Alternative Housing Supply Elasticities (b) Additional Redevelopment of Public Housing



Notes: This figure displays the percent change in rents due to demolitions under alternative housing policies. Panel A reports results from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Our baseline results are based on a housing supply elasticity of 0.163. Panel B reports results from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. Our baseline results are given by 0 percent additional redevelopment.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

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. Consistent with our neighborhood-level analysis in Section 8.2.2, 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 15.6 percent in our baseline specification when the housing supply elasticity is assumed to be 0.163. Increasing the housing supply elasticity to

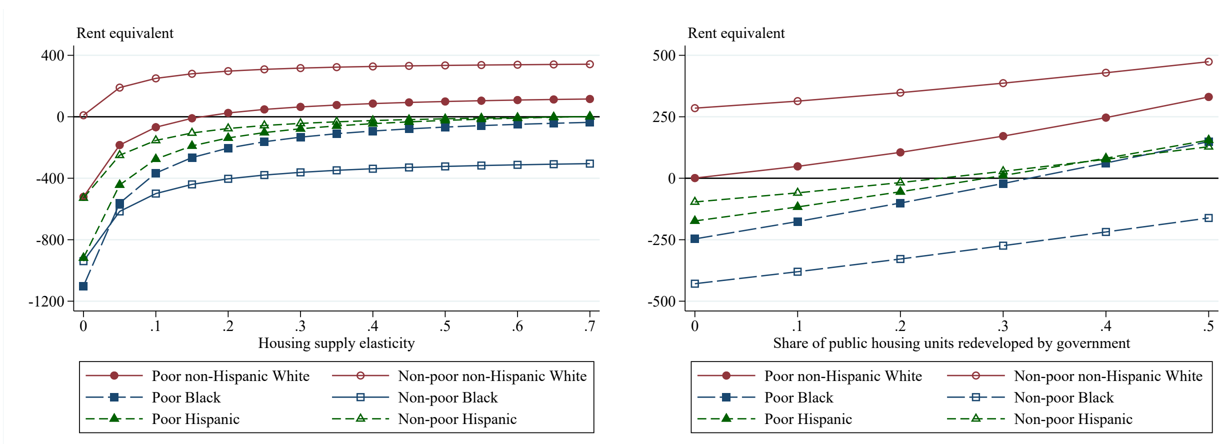
0.45—approximately equal to the average elasticity for Youngstown, Ohio, or Gary, Indiana, estimated in Baum-Snow and Han (2024)—reduces the effect on rents by more than half to 6.7 percent. The blue line (square marker) shows that the qualitative patterns in the neighborhoods indirectly impacted are similar but muted relative to the neighborhoods where demolitions occur.

Panel A of Figure 12 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 -\$247 to -\$78 in terms of rent equivalent welfare 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.[48] 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 (2024).

Figure 12: Consequences of Public Housing Demolitions on Welfare Under Alternative Housing Policies

(a) Alternative Housing Supply Elasticities

(b) Additional Redevelopment of Public Housing



Notes: This figure displays the rent equivalent welfare effect of public housing demolitions under alternative housing policies. Panel A reports results from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Panel B reports results from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

9.2 Welfare Effects under Housing Redevelopment

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 demol-

[48] Appendix Figure A.6 provides welfare estimates separately for renters and homeowners.

ished. 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 neighborhoods (see Appendix Table A.1).[49] 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, we report how the demolition effects on rents and welfare 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 percent of the units that we observe as having been destroyed by 2010. This upper bound is motivated by the fact that about 50 percent of former public housing sites were not re-developed for residential or commercial uses (Appendix Table A.1).

In Panel B of Figure 11, 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 20 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: 1.6 percent), although this change is 25 percent lower than the increase without additional redevelopment.

The results in Panel B of Figure 12 show that the impact of demolition on welfare improves for most groups with the scale of redevelopment. For all but non-poor Black households, redeveloping 35 percent of destroyed housing is a sufficient intervention that results in welfare gains associated with demolition. In all cases, however, White households benefit by more than Black and Hispanic households.

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 more beneficial 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 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

[49] 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.

implications by indirectly targeting minorities through their location choices.

10 Welfare Impacts on Public Housing Residents

Our primary approach using Census data facilitates a comprehensive assessment of the welfare consequences of public housing demolitions across many demographic and economic groups. However, a limitation of this approach is that it does not isolate the welfare consequences for initial public housing residents. While displaced public housing residents comprise a small share of the total population in Chicago, they are policy-relevant because they were forced to move by the HOPE VI demolitions. In this section, we directly study how public housing demolition shaped welfare for former public housing residents using detailed administrative data.

10.1 Model

Our modeling approach captures two unique aspects of location choice for households who were displaced from public housing. First, these individuals lost access to in-kind public housing benefits. Instead, they received Section 8 housing vouchers that could be used to rent housing from the private market.[50] Importantly, the rent contribution for public housing and the voucher have similar program and rent rules. As a result, the transition from public housing to vouchers should not mechanically affect the income of displaced households.[51] Second, these individuals incurred the utility costs of being displaced.

We account for these distinct aspects of the neighborhood choice problem by augmenting our baseline model of neighborhood demand. In particular, the indirect utility of household i whose public housing residence was in neighborhood j_0 is:

$$\begin{aligned} V_{ij} = \alpha_p \ln(p_{ij}) + & \overbrace{\alpha_b b_j + \alpha_h h_j + \alpha_{Inc} \ln(Inc_j) + \alpha_{PH} PH_j + \theta x_j + \xi_j}^{\delta_j} \\ & - MC_{\text{unit}} \cdot \mathbb{1}\{\text{change unit}\} - MC_{\text{nhood}} \cdot \mathbb{1}\{j(i) \neq j_0\} - MC_{\text{county}} \cdot \mathbb{1}\{\text{change county}\} + \epsilon_{ij}. \end{aligned} \tag{12}$$

This expression differs from the indirect utility function in equation (1) in several notable ways. First, the price of housing, p_{ij}, varies across individuals in the following way. If a public housing resident was not displaced by demolition, denoted by $D_i = 0$, they must pay a standard cost-of-living amount c_j when living in public housing.[52] If individuals are not displaced from public housing and choose to leave public housing, then they must also pay the full market rent, r_j and cost of living c_j. Finally, if individuals are displaced from public housing, denoted by $D_i = 1$, then the price they face is $p_{ij} = \max\{c_j, c_j + r_j - v\}$, where v is the amount of the Section 8 voucher.[53] The second key difference is that we incorporate three types of moving costs. Individuals pay a moving cost, MC_{unit} when they move out of their public housing unit. In

[50] The voucher subsidy program has a maximum benefit which was equal to the difference between the gross unit rental price or the local Fair Market Rent (FMR) and the family's required rental contribution. The FMR is specified based on the number of bedrooms and during our time period was equal to the fortieth percentile of rent for the Chicago-area private market distribution.

[51] In both the public housing or voucher assistance programs, assisted households must pay 30 percent of their adjusted income as a rental contribution.

[52] In our empirical analysis, we calculate c_j as the difference between median gross rent (which includes utilities) and median contract rent (which does not).

[53] We compute subsidies using the history of 2-bedroom fair market rents provided by the HUD for Cook County. The subsidy was \$778.23 and \$1036.62 in 2017 dollars for 2000 and 2010 respectively.

addition, they pay a moving cost equal to MC_{nhood} and MC_{county} if they move to a new Census tract within Cook County or out of Cook County, respectively.[54] The remaining elements of indirect utility, δ_j and ϵ_{ij}, are identical to our baseline model.

10.2 Data, Estimation, and Identification

We estimate neighborhood preferences for former public housing residents using a sample of public housing residents created by combining building records from the CHA and social assistance case files from the Illinois Department of Human Services (IDHS). The building records identify addresses of each public housing building that was subject to demolition while the IDHS records allow us to identify individuals and track them over time while they remain on social assistance.[55] These data cover 2,996 households who resided in public housing that was and was not selected for demolition between 1995 and 1998.[56] The residents of the non-demolished buildings can be used as a comparison group for identifying parameters of our location choice model.[57] For each household, we observe whether they continue to live in public housing or, if they move, their first post-public-housing residential address. We combine these administrative records from PHA with Census data used in our main analysis.

We use a two-step approach that combines maximum likelihood estimation (MLE) and our baseline instrumental variable and difference-in-difference approach. In our first step, we use MLE to estimate $(\alpha_p, \delta_j, MC_{\text{nhood}}, MC_{\text{unit}})$ using the individual level data on displaced and non-displaced public housing households. We follow Galiani, Murphy and Pantano (2015), who combine MLE with random variation from the Moving to Opportunity experiment, in maximizing a log likelihood function subject to the constraint that choice probabilities from the model match the empirical choice probabilities, $\tilde{P}_{jk}$, for low-income Black households in the 2000 Census. Appendix C has additional details. In the second step, we use 2SLS to regress the estimated common component of utility, $\hat{\delta}_j$, on the underlying structural factors:

$$\hat{\delta}_j = \alpha_b b_j + \alpha_h h_j + \alpha_{Inc} \ln(Inc_j) + \alpha_{PH} PH_j + \theta x_j + \xi_j. \tag{13}$$

We treat the vector $(b_j, h_j, \ln(Inc_j))$ as endogenous and use the same instruments as in our baseline approach.

The identification of the model parameters is enhanced by our use of individual-level variation in exposure to public housing demolitions. Chyn (2018) shows there are no detectable differences in the baseline characteristics of displaced and non-displaced residents. Idiosyncratic maintenance problems, such as flooding due to the bursting of building pipes, appear to drive a substantial amount of the variation in which buildings were demolished during the early demolition period. In our setting, exogenous assignment of displacement status, D_i, would lead to exogenous differences in the price of housing, p_{ij}. Intuitively, the cost of housing in a different neighborhood is lower for residents of a demolished building because they gain

[54]Specifically, individuals moving to another neighborhood in Cook County pay $MC_{\text{unit}} + MC_{\text{nhood}}$. Individuals who move out of Cook County pay $MC_{\text{unit}} + MC_{\text{nhood}} + MC_{\text{county}}$.

[55]Chyn (2018) shows that there are no detectable impacts of demolitions on whether former public housing residents have social assistance.

[56]Data for individuals who resided in public housing in later years is unfortunately not available.

[57]The buildings that were not selected for demolition in this analysis remained open for many years after the initial wave of demolitions in Chicago.

access to Section 8 vouchers, unlike the individuals whose building is not demolished. This individual-level variation is used to identify the sensitivity of individuals to housing prices, α_p. A key assumption for this approach is that neighborhood-level unobservables, ξ_j, are the same for displaced and non-displaced individuals. The identification of the remaining parameters follows a similar argument as in our main model as presented in Section 7.2.

10.3 Preference and Welfare Results

Table 3 reports estimates of the preference parameters in equation (12). The results differ from preference estimates from Section 7.3 along several dimensions. First, households initially living in public housing, who are especially poor, are more sensitive to housing costs than what we observe for the broader set of poor Black households. Households that initially live in public housing also have a strong preference for living in neighborhoods with public housing—a finding that is consistent with the possibility that they value social networks based in public housing. Finally, these households face substantial utility costs from moving. In particular, the estimates imply that the ten-year utility cost of moving out of public housing is equivalent to \$25,121, the cost of moving to a different tract is \$44,959, and the cost of moving out of Cook County is \$27,738.[58] These moving cost estimates are in line with those from previous work.[59]

Next, we quantify the welfare effects of demolitions on displaced public housing residents sequentially in a series of exercises that vary the individual and housing market conditions faced by households. An important caveat for the interpretation of our results is that we exclude direct consideration of the effects of relocating on health and economic outcomes for displaced individuals. Prior research has documented that moves from disadvantaged public housing projects into to lower poverty areas generate notable improvements on adult health and long-run labor market outcomes of children (Kling, Liebman and Katz, 2007; Chetty, Hendren and Katz, 2016; Chyn, 2018).

The first four rows of Table 4 quantify the same channels that we consider for the broader sample of non-public-housing renters in Table 2. We proceed by calculating the associated changes in public housing resident sorting and welfare after sequentially simulating scenarios where we remove public housing as a neighborhood characteristic in each tract, allow the overall Black and Hispanic population shares to adjust in response to the removal of public housing and the associated changes in demographics, and impose equilibrium price adjustments.[60] Taken together, these channels lower the welfare of public housing residents by an amount equal to a \$2,779 increase in annual rents. This is a larger decrease in welfare than what is experienced by the broader sample of renters studied in Table 2. The key explanation is that public housing residents view the intensity of public housing as a neighborhood amenity.

[58] We treat the moves observed in the administrative data as occurring over a one-year period for the purpose of estimating the model. Thus, the estimates in Table 3 are measures of one-year moving costs. Our counterfactual analysis using decadal data assumes that the pre-demolition and post-demolition periods are ten years apart. We convert annual moving costs to a ten-year moving cost using a yearly discount factor of 0.95. We estimate a marginal utility of income for public housing residents equal to 6.35e-5.

[59] As noted above in Section 8.3, the estimates in Bayer et al. (2016) imply a one-year moving cost for Chicago that is approximately equal to \$38,000, and a ten-year moving cost of \$22,752.

[60] In the simulations associated with the first four rows, we do not force public housing residents to move out of public housing. Instead, we vary the associated neighborhood characteristics, which changes utility, and allow public housing residents to move if they find it optimal to do so.

Table 3: Estimates of Neighborhood Preference Parameters for Public Housing Sample

	(1)
Log median rent	-0.378***
	(0.052)
Black population share	0.513***
	(0.082)
Hispanic population share	-0.550***
	(0.155)
Log median household income	-0.452***
	(0.113)
Public housing units as a share of housing stock	0.913**
	(0.396)
Moving cost: Leaving public housing unit	-2.666***
	(0.137)
Moving cost: Leaving tract	-4.772***
	(0.095)
Moving cost: Leaving Cook County	-2.944***
	(0.229)
Specifications include:	
Tract Fixed Effects	✓
Log median number of rooms	✓
Log median year built	✓
Homeownership share	✓
Land use variables	✓
Observations (households)	2,996

Notes: This table presents regression results of preference parameters for the model of initial public housing residents described in Section 10.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

In rows 5 through 7 of Table 4, we expand on these results by considering the unique conditions faced by public housing residents. First, households experience the disutility of moving out of their housing unit. The average welfare loss, measured in rent equivalent units, grows to $27,901. This is due to the large estimated utility cost of moving out of public housing. A distinctive feature of our setting is that public housing residents are forcibly displaced, which means that they incur this large moving cost even if they do not receive a particularly attractive idiosyncratic preference shock for another neighborhood. Next, households are forced to pay market rents, which lowers their welfare by $2,415. Finally, households receive a housing voucher that can be used anywhere in the city; this raises their welfare by $10,178. The magnitude of this increase in welfare implies that there are large gains that stem from the fact that households have more neighborhood choice when they receive a voucher. Overall, we estimate that demolitions reduced the welfare of displaced public housing households by $20,137. The key forces that affect the welfare of displaced public housing residents are the utility costs of being displaced and the benefit of gaining access to a Section 8 voucher.

Can alternative housing policies improve the welfare of public housing residents? Appendix Figure A.8 reports the welfare effects of demolitions under the same counterfactual housing policies that we considered for the broader sample. The improvements in welfare from increasing the housing supply elasticity or extent of redevelopment are similar for public housing residents as for the broader set of poor Black renters. Nonetheless, public housing residents continue to experience meaningful welfare losses, primarily because

Table 4: The Welfare Effects of Public Housing Demolitions for Households Originally Living in Public Housing

	Change from baseline (no demolitions)
Destroy buildings in tract	-2,285.6
... and change neighborhood composition via demolitions	-2,670.7
... and change neighborhood composition via resorting	-2,653.9
... and change housing prices (all channels faced by non-public-housing renters)	-2,779.1
... and forced to move out of housing unit	-27,900.5
... and forced to pay market rents	-30,315.4
... and provided with housing voucher (all channels)	-20,137.2

Notes: This table reports the rent equivalent change in welfare for each counterfactual compared to a counterfactual with no public housing demolitions for households that originally live in public housing. A positive rent equivalent implies that households are better off in the indicated counterfactual relative to the counterfactual with no public housing demolitions. The first four rows consider the same counterfactual scenarios as for non-public-housing renters in the first four rows of Table 2, Panel A, while the last three rows consider situations that are unique to public housing residents. In the first row, we consider a counterfactual in which public housing is removed as a neighborhood characteristic 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 non-public-housing-residents re-optimize their location choices and displaced public housing residents seek market-based housing. The fourth row allows housing prices to adjust in addition. In the fifth row, households experience the disutility of moving out of their housing unit. In the sixth and seventh rows, households must pay market rents and receive housing vouchers, respectively.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

of moving costs arising from forced displacement.

11 Aggregate Consequences of Public Housing Demolition

As a final exercise, we provide a simple assessment of the welfare consequences of public housing demolitions on all households living in Chicago, the general population captured in our Census data and the set of households who were directly displaced from public housing. While this quantification has the potential to shed light on the overall welfare effects of demolitions, we stress caution when interpreting these results due to distinctions in the neighborhood choice models that we estimate for the general population and public housing residents. For example, our modeling of moving costs in our model of public housing residents is internally estimated and includes multiple types of moving costs for additional flexibility. In contrast, we rely on an external estimate from Bayer et al. (2016) to incorporate moving costs into welfare analysis for the general population of households.

Table 5 reports both average and aggregate welfare consequences for each group that we consider in our analysis. The aggregate welfare effect is equal to the average welfare effect times the number of households in column 1. Columns 2 and 4 report results without moving costs, while columns 3 and 5 include them. Without imposing moving costs, we estimate that demolitions led to an aggregate increase in welfare of around $100 million, with large gains for both non-poor White households and households displaced from public housing, and large losses for other Black households. After incorporating moving costs, we find that demolitions reduced welfare by around $380 million with losses for those displaced from public housing and minority households.[61] As noted above, the large negative impact for public housing households is

[61] Panel A of Appendix Figure A.4 examines the sensitivity of the estimated average welfare impacts to moving costs. The welfare results for the census sample are fairly robust to the assumed value of moving costs, as noted above in Section 8.3, as moves happen infrequently.

Table 5: The Aggregate Welfare Effects of Public Housing Demolitions

	Total households	Average welfare change		Aggregate welfare change, millions	
	(1)	(2)	(3)	(4)	(5)
Households not starting in public housing					
Non-Hispanic White, Poor	120,840	0.9	-36.4	0.1	-4.4
Non-Hispanic White, Non-poor	786,279	285.1	353.5	224.1	277.9
Black, Poor	142,858	-246.6	-402.1	-35.2	-57.4
Black, Non-poor	300,190	-428.9	-808.8	-128.7	-242.8
Hispanic, Poor	57,177	-173.6	-296.2	-9.9	-16.9
Hispanic, Non-poor	254,458	-96.4	-202.5	-24.5	-51.5
Households starting in public housing	13,985	5,328.8	-20,137.2	74.5	-281.6
Total	1,675,787	59.9	-224.8	100.4	-376.8
Welfare effects include moving costs		No	Yes	No	Yes

Notes: This table reports the aggregate 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 due to demolitions relative to the counterfactual with no public housing demolitions. Columns 2 and 3 report the average welfare change for households of the indicated type, and columns 4 and 5 report the aggregate welfare change in millions of dollars, which is found by multiplying the average changes in columns 2 and 3 by the total number of households in column 1 and then converting to millions. Columns 2 and 4 report welfare calculations that assume there are no moving costs, while columns 3 and 5 report welfare calculations that are based on calibrated moving costs for households not starting in public housing and estimated moving costs for households starting in public housing. Statistics on total households by group in Cook County are based on the 2010 Census.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

driven by moving costs associated with forced displacement. Moving costs quantitatively matter less in the general population since the small fraction of households that choose to move do so only when their moving costs are offset by the gains to relocation.

There are natural political economy questions in light of our results suggesting that aggregate welfare declined when moving costs are taken into account. Non-poor White residents constitute less than half of the population of Chicago and their welfare increases by $278 million due to demolitions. While this group falls short of the simple majority needed to win an election, an important consideration is the fact non-poor White households are more likely to be registered voters and turn out during elections (U.S. Census Bureau, 2010). Hence, local policymakers in Chicago have a strong electoral incentive to promote policies that may disproportionately benefit their core political constituents.

There is more nuance when considering the perspective of a social planner deciding whether to promote demolitions that have heterogeneous impacts across groups. Given the aggregate decline in welfare that we document, a planner could only support demolitions if non-poor White households have an outsized weight in their social welfare function. For instance, a planner would only support demolitions if they value the $278 million gains for non-Poor White households 2.3 times higher to outweigh the $639 million welfare loss for all remaining groups.

12 Conclusion

This paper provides new evidence on the welfare consequences of neighborhood redevelopment programs by studying federally-funded public housing demolitions in Chicago. As noted by prior studies, 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 non-poor White households alongside welfare losses for minority households. The unequal effects of demolitions arise from two important forces. First, while most households benefit from the destruction of public housing, reductions in the racial minority share in targeted neighborhoods and subsequent re-sorting 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.

We also present supplemental results from an analysis of displaced public housing residents that show large negative impacts of demolition that are driven by moving costs that must be absorbed due to forced relocation. These moving costs offset the welfare benefits that displaced households gain from the fact that housing vouchers expand neighborhood choice. The total negative impact on public housing residents outweighs the gains that non-poor White households—the largest demographic group in Chicago—experience in our setting.

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. The findings in this paper should be relevant in other settings where housing policies generate large-scale re-sorting and preferences over racial composition and price sensitivity are similar to those in our context.

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, Shamsuddin and Kelly, 2018) may have muted the welfare benefits of public housing demolition for minority and lower-income residents.

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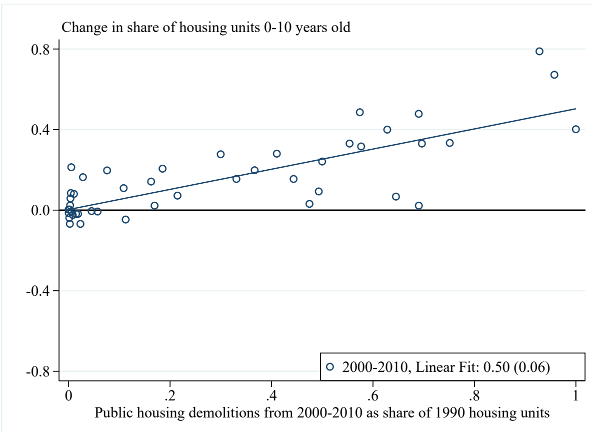
Online Appendices

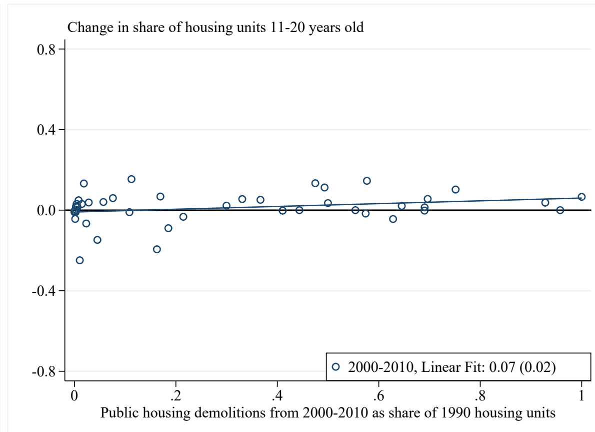
A Appendix Figures and Tables

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

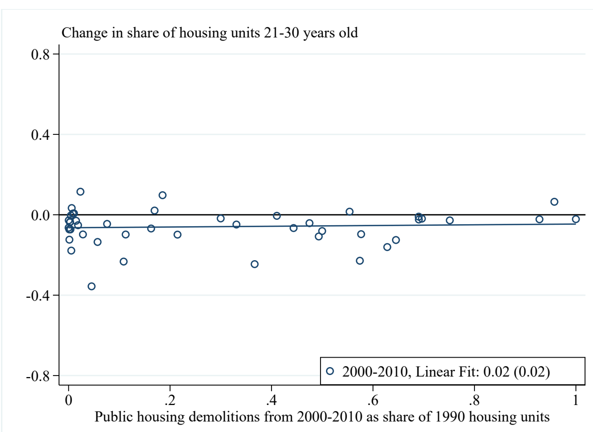
(a) Share of Housing Units 0–10 Years Old



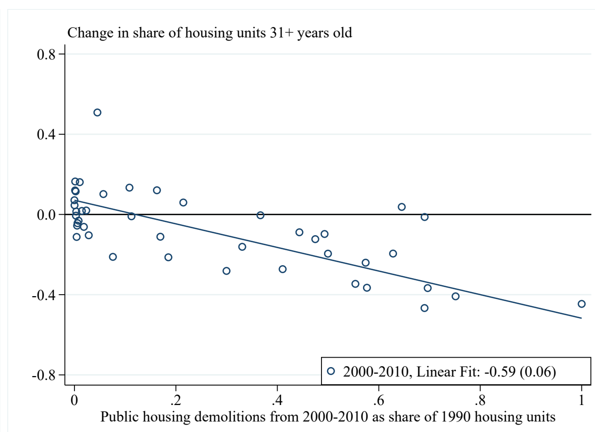
(b) Share of Housing Units 11-20 Years Old



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



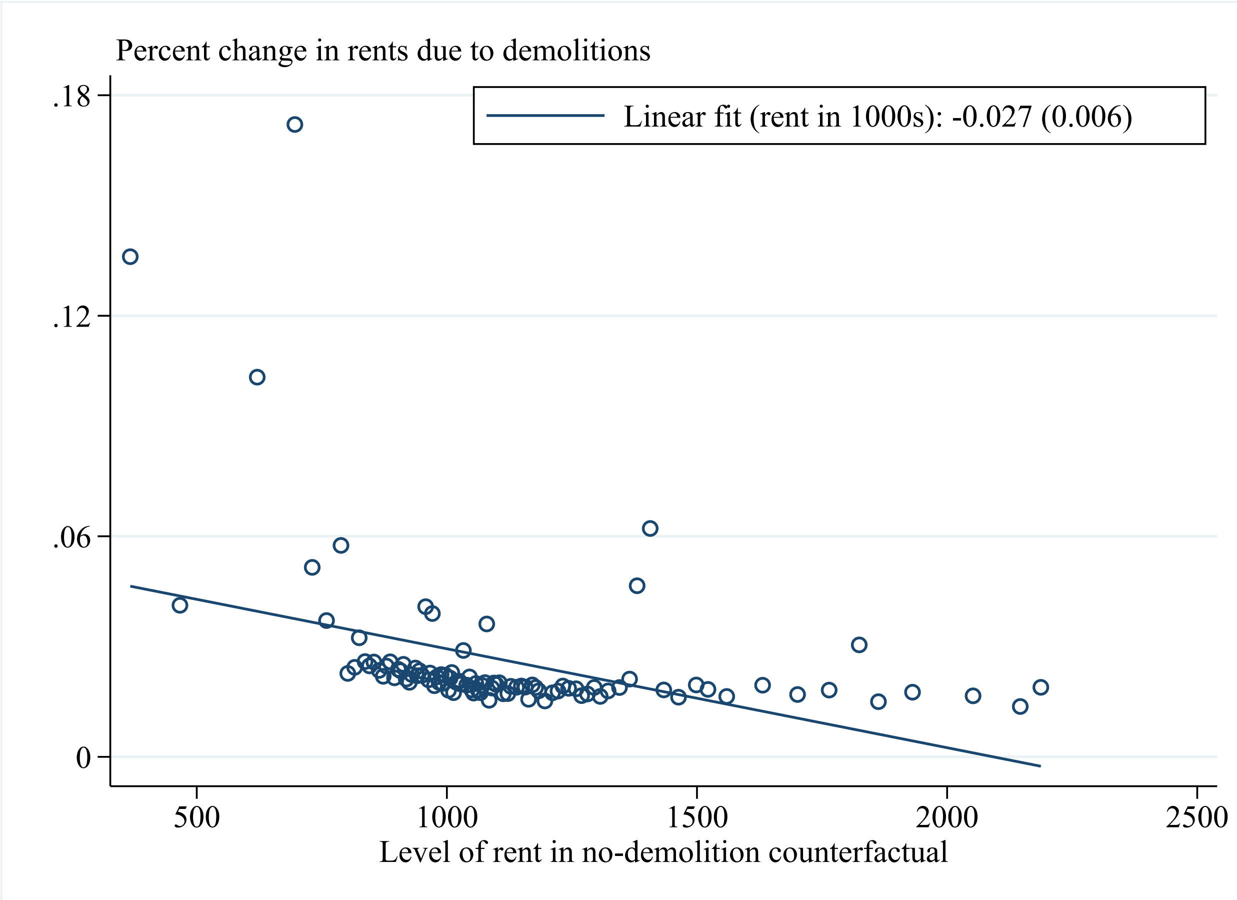
(d) Share of Housing Units 31+ Years Old



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.

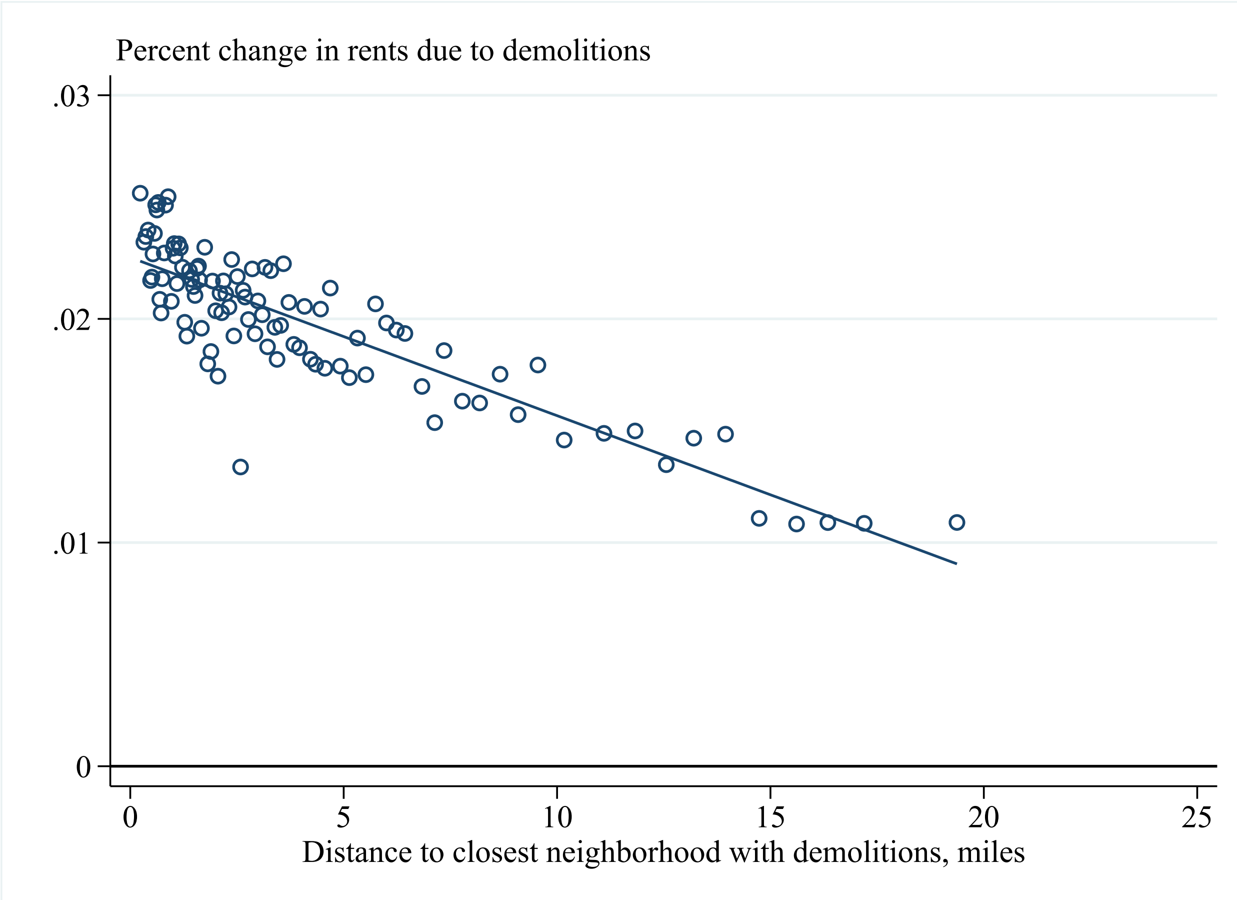
Appendix Figure A.2: Tract-Level Rent Changes Due to Public Housing Demolitions Relative to Rents in Absence of Demolitions



Notes: This figure displays the percent change in median 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 thousands of dollars. The bin scatter is constructed for 100 percentiles.

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

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

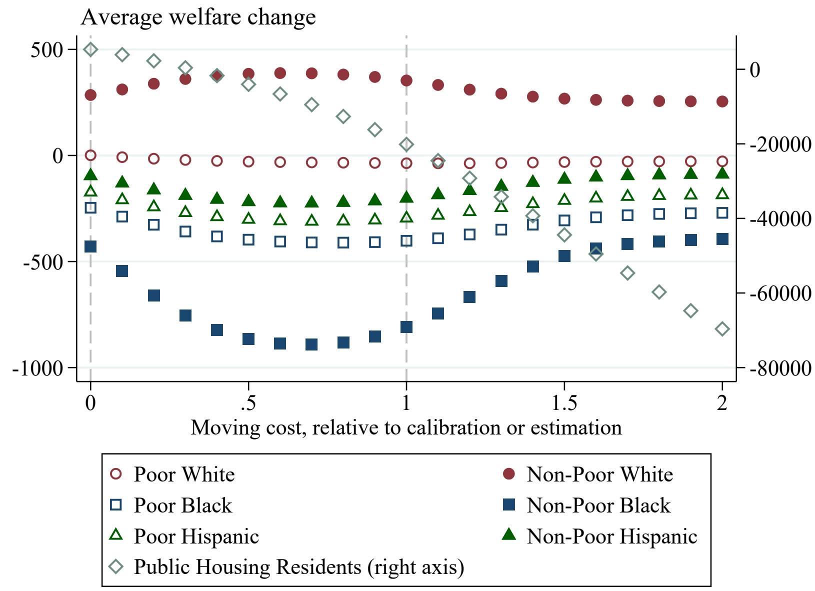


Notes: Figure displays the percent change in median rents due to public housing demolitions for tracts that did not have public housing against the distance to the closest tract with demolitions. We construct the dependent variable using estimates from the model and a comparison of differences between 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.

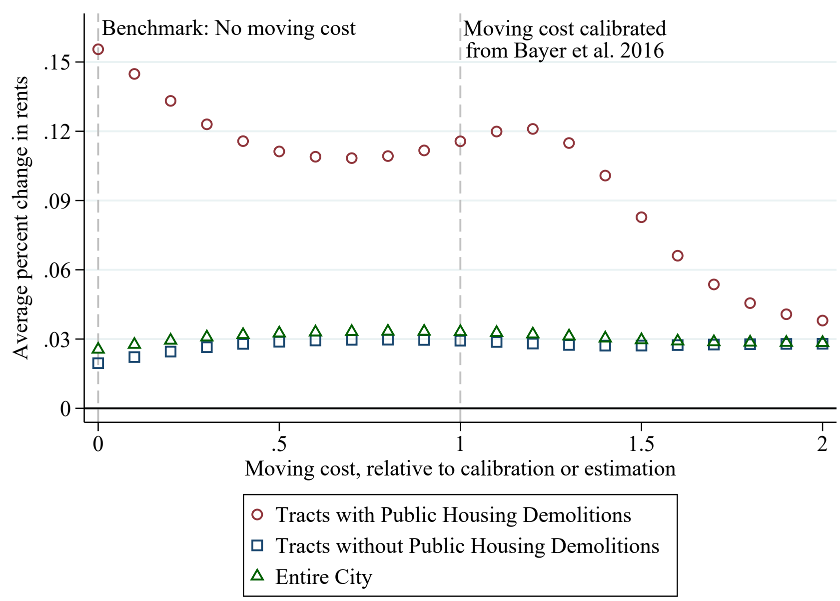
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.4: Assessing Sensitivity of Results to Different Assumed Moving Costs

(a) Average Welfare



(b) Rent

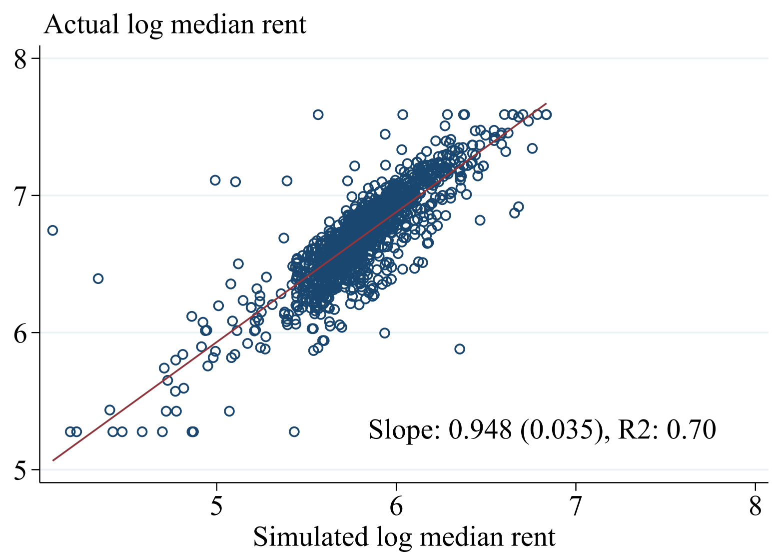


Notes: Panel A plots the rent equivalent change in average welfare compared to a counterfactual with no public housing demolitions under different assumed values of a one-year moving cost. A positive rent equivalent implies that households are better off in the situation with demolitions relative to the counterfactual with no public housing demolitions. Our benchmark model assumes no moving costs, as indicated in the left-most numbers. We calibrate a one-year moving cost of $38,000 based on the estimates of Bayer et al. (2016) and summary statistics on Chicago. Panel B plots the average percent change in rents due to demolitions for tracts with a public housing demolition, tracts without a public housing demolition, and all of Cook County, IL.

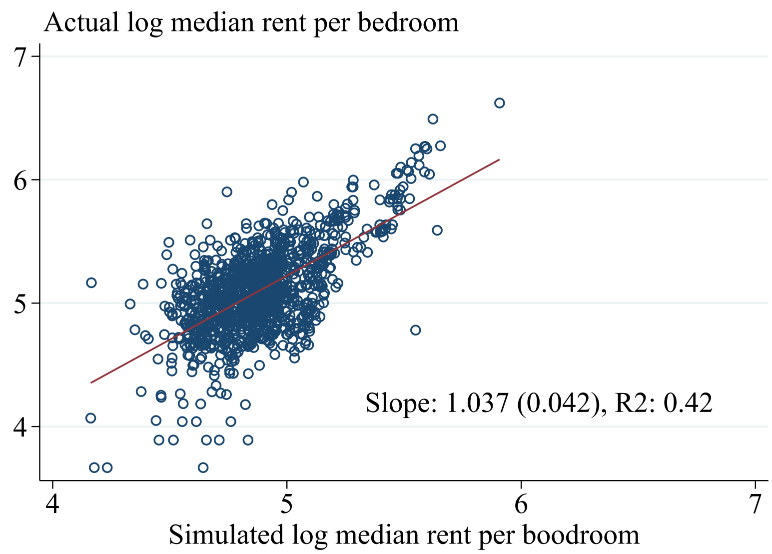
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.5: Comparing Out-of-Sample Fit of Baseline Model and Cobb-Douglas Model Using Rent Data

(a) Baseline Model



(b) Cobb-Douglas Model

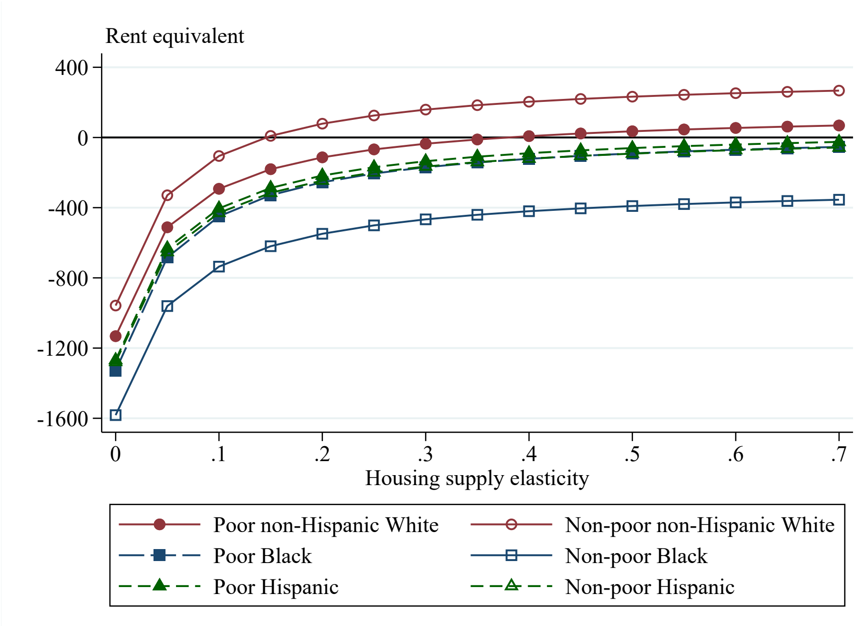


Notes: Panel A repeats the plot from Figure 7 of actual log rents in 1990 in census tracts against log rents that are simulated by an out-of-sample procedure. In particular, we construct simulated rents for 1990 using the coefficients and tract fixed effects estimated using 2000–2010 data, exogenous 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). Panel B presents the actual log median rent minus the actual log median number of bedrooms in census tracts against the log rent implied by the model estimates minus the actual log median number of bedrooms. We follow the same procedure as in Panel A, but using the Cobb-Douglas model.

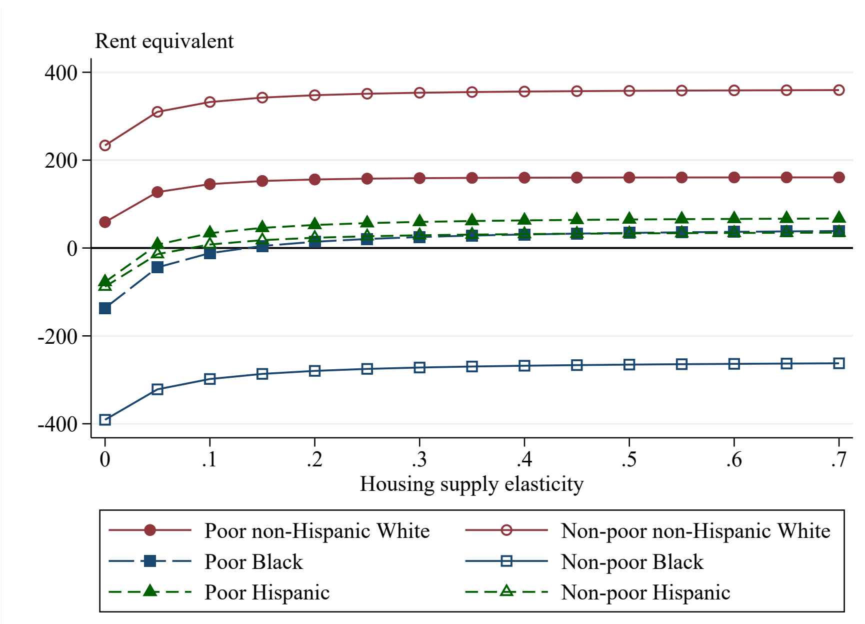
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.6: 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

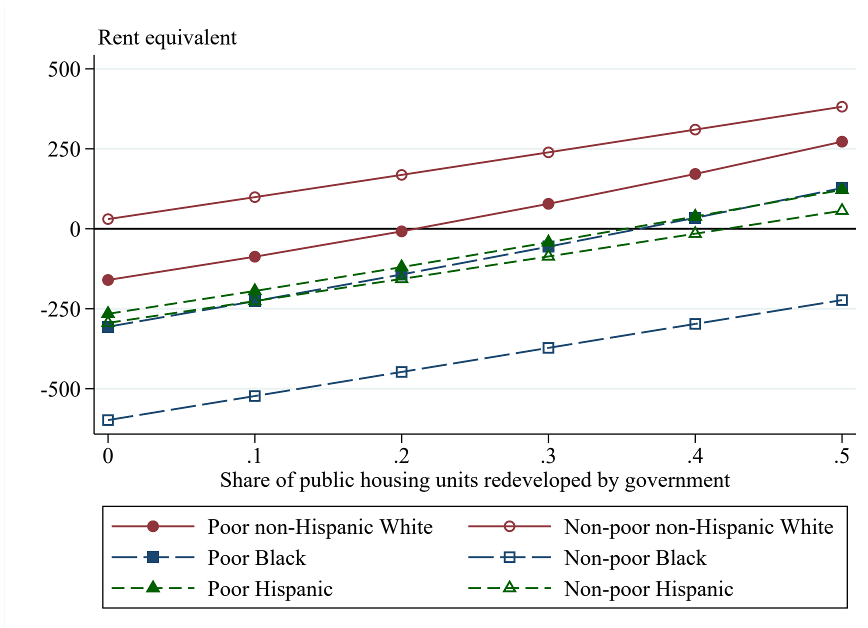


Notes: Figure displays outcomes from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Panel A shows the rent equivalent welfare effect of public housing demolitions for renters and Panel B shows analogous results for homeowners. Our baseline results are based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

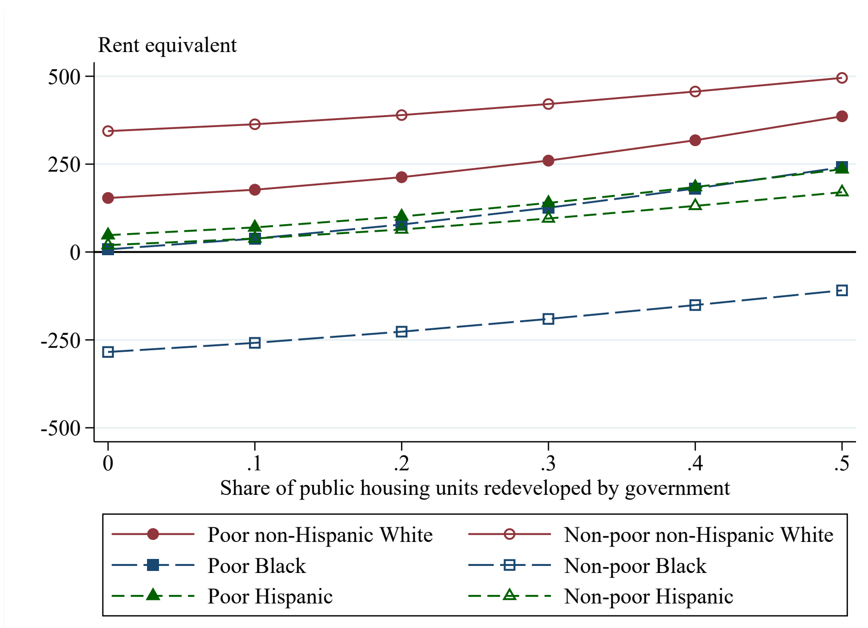
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.7: 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

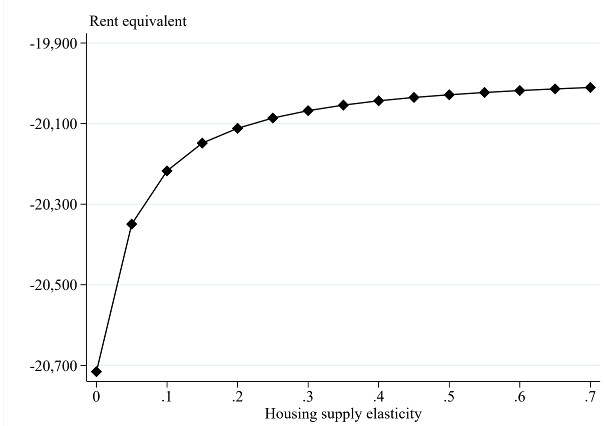


Notes: Figure displays outcomes from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. Panel A shows the rent equivalent welfare effect of public housing demolitions for renters and Panel B shows analogous results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

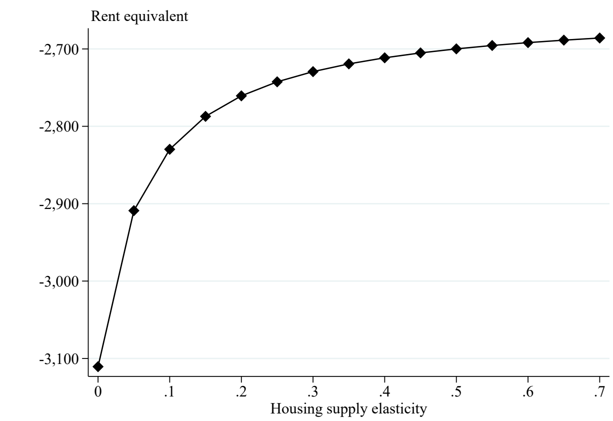
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.8: Consequences of Public Housing Demolitions on Welfare under Alternative Housing Supply Elasticities and Additional Redevelopment of Public Housing for Households in Public Housing

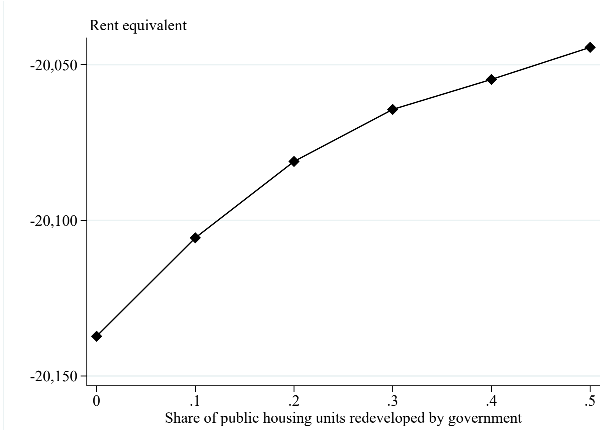
(a) Alternative Housing Supply Elasticities, All Channels



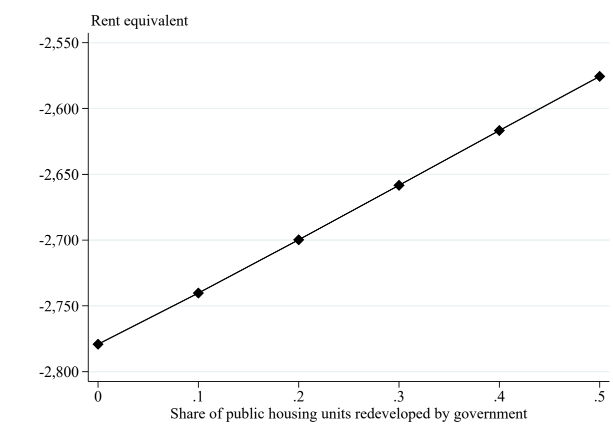
(b) Alternative Housing Supply Elasticities, Channels Faced by Non-Public-Housing Renters



(c) Additional Redevelopment of Public Housing, All Channels



(d) Additional Redevelopment of Public Housing, Channels Faced by Non-Public-Housing Renters



Notes: Panels A and B display outcomes from counterfactual scenarios in which the housing supply elasticity takes on the indicated value. Panels C and D display outcomes from counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. All panels show the rent equivalent welfare effect of public housing demolitions for individuals who live in public housing before demolitions take place. Panels A and C incorporate all channels faced by public housing residents, while Panels B and D only focus on the channels faced by non-public-housing renters, as shown in Table 4. Our baseline results are given by a housing supply elasticity of 0.163 (Panel A) and 0 percent additional redevelopment (Panel B).

Source: Authors' calculations using data from the Chicago Housing Authority and U.S. 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, Ferreira and McMillan (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, Ferreira and McMillan (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 percent of tracts and 90 percent 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.

Appendix Table A.8: 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 percent change in rents	0.156	0.020	0.025
Share of total rent increase	0.28	0.72	1.000

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 provides 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.

Appendix Table A.9: The Welfare Effects of Public Housing Demolitions, Robustness to Accounting for Moving Costs

	Change from baseline (2010 Census)					
	Non-Hispanic White		Black		Hispanic	
Counterfactual scenario	Poor (1)	Non-poor (2)	Poor (3)	Non-poor (4)	Poor (5)	Non-poor (6)
Panel A. Results without accounting for moving costs						
All channels for renters	-160	30	-306	-598	-266	-294
All channels for owners	154	344	8	-284	48	20
Average welfare change across renters & owners	1	285	-247	-429	-174	-96
Panel B. Results accounting for moving costs						
All channels for renters, besides moving costs	221	-22	332	709	300	340
All channels for renters	-251	14	-482	-1,035	-419	-466
All channels for owners	167	432	-63	-616	-1	-48
Average welfare change across renters & owners	-36	353	-402	-809	-296	-203

Notes: This table reports the rent equivalent change in welfare compared to a counterfactual with no public housing demolitions. A positive rent equivalent implies that households are better off in the situation with demolitions relative to the counterfactual with no public housing demolitions. Panel A reports the overall welfare changes from Table 2 for our baseline model, which does not include moving costs. Panel B reports welfare changes from the model described in Section 8.3.1, which does account for moving costs. The first row of Panel B excludes the welfare costs associated with paying moving costs, which allows us to gauge the importance of moving costs by comparing rows 1 and 2.

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

Appendix Table A.10: Comparison of Willingness to Pay Parameters

	Poor Households			Non-Poor Households		
	Non-Hispanic White	Black	Hispanic	Non-Hispanic White	Black	Hispanic
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Baseline Model						
Black population share	44	-134	-135	294	-710	-210
Hispanic population share	36	-1	-152	312	-286	-306
Log median household income	-24	-18	-6	-47	-32	4
PH units as a share of housing stock	122	123	135	139	285	152
Panel B: Cobb-Douglas Model						
Black population share	35	-67	-62	86	-106	-86
Hispanic population share	26	14	-64	75	-23	-120
Log median household income	-24	-19	-15	-32	-26	-13
PH units as a share of housing stock	123	130	135	131	158	142

Notes: This table reports the estimated willingness to pay for a 1 percentage point increase in the Black population share, Hispanic population, or share of the housing stock that is public housing or a 1 percent increase in the median household income of a neighborhood's residents. Panel A presents results for our baseline model. Panel B presents results for a Cobb-Douglas model.

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:

$$\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha) = \mathcal{S}_j(p_j) \qquad \forall j = 1, ..., J \tag{B.1}$$
$$\frac{\mathcal{D}_j^B(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha)}{\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha)} = b_j \qquad \forall j = 1, ..., J \tag{B.2}$$
$$\frac{\mathcal{D}_j^H(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha))}{\mathcal{D}_j(\mathbf{p}, \mathbf{b}, \mathbf{h}, \mathbf{x}, \xi; \alpha))} = h_j \qquad \forall j = 1, ..., J. \tag{B.3}$$

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:

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

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

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 demographics act as an agglomeration force.[62]

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

[62]For numerical stability, we recommend dividing $\mathcal{EDH}^n$ by $\mathcal{D}_j(\mathbf{p}^n, \mathbf{b}^n, \mathbf{h}^n) + \mathcal{S}_j(\mathbf{p}^n)$ in Step 2 of the algorithm.

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.

C Details on Estimation of Model for Public Housing Residents

Equation (12) describes how households choose their optimal location. To simplify notation, we rewrite this as

$$V_{ij} = \delta_j + \alpha_p \ln(p_{ij}) - MC_{ij} + \epsilon_{ij}, \tag{C.1}$$

where

$$\delta_j \equiv \alpha_b b_j + \alpha_h h_j + \alpha_{Inc} \ln(Inc_j) + \alpha_{PH} PH_j + \theta x_j + \xi_j$$

describes the utility component common to all households for a given choice j, and

$$MC_{ij} \equiv MC_{\text{unit}} \cdot \mathbb{1}\{\text{change unit}\} + MC_{\text{nhood}} \cdot \mathbb{1}\{j(i) \neq j_0\} + MC_{\text{county}} \cdot \mathbb{1}\{\text{change county}\}.$$

describes all of the moving costs incurred by the household for a given choice. We denote the vector of moving costs by $MC \equiv (MC_{\text{unit}}, MC_{\text{nhood}}, MC_{\text{county}})$.

The probability that household i chooses neighborhood j is given by:

$$\mathcal{P}^i_j = \frac{\exp(\delta_j + \alpha_p \ln(p_{ij}) - MC_{ij})}{\sum_{j'} \exp(\delta_{j'} + \alpha_p \ln(p_{ij'}) - MC_{ij'})}, \tag{C.2}$$

and the contribution of household i to the likelihood function is:

$$\prod_i (\mathcal{P}^i_j)^{y_{ij}},$$

where $y_{ij} = 1$ if household i chooses to live in neighborhood j, and 0 otherwise.

Given this set-up, we can write the log-likelihood function for the sample of public housing residents as

$$\mathcal{L}(\delta_j, \alpha_p, MC) = \sum_i \sum_j y_{ij} \ln \mathcal{P}^i_j. \tag{C.3}$$

As in Galiani, Murphy and Pantano (2015), we constrain the maximization of the log-likelihood function by requiring that model-implied choice probabilities for each neighborhood are equal to the empirical choice probabilities, $\tilde{P}_{jk}$, for low-income Black households in the 2000 Census.[63] Our motivation for imposing these constraints is identical to Galiani, Murphy and Pantano (2015). In particular, the CHA data that we use to estimate the location choice model has too few observations and many locations do not receive any individual in the sample. Because our estimation procedure, also requires estimates δ_j for all j in a first step, the unconstrained problem would set $\delta = -\infty$ for those j that do not receive any public housing residents, which creates a numerical problem. Imposing the constraints mentioned above helps overcoming this identification problem.

[63] Nearly all residents of Chicago's public housing were Black in the time period that we study.