The Welfare Impacts of Urban Renewal and Forced Relocation: Evidence from Chicago's HOPE VI Demolitions*

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Abstract

This paper studies one of the largest spatially-targeted redevelopment efforts in the United States: public housing demolitions sponsored by the HOPE VI program. Focusing on Chicago, we estimate a structural model of neighborhood choice that combines administrative records tracking displaced public-housing residents with Census/ACS data capturing citywide sorting and prices. In the administrative records, displaced residents' moves are typically incremental and rarely reach highly advantaged neighborhoods. Our structural estimates show large welfare losses driven primarily by displacement costs, which voucher-enabled choice only partly offsets. Demolitions also raise equilibrium housing prices and reshape neighborhood composition, yielding welfare gains concentrated among higher-income homeowners. Counterfactual simulations highlight a sharp contrast: targeted demand-side policies—such as providing more generous housing vouchers or intensive counseling—mitigate welfare losses experienced by displaced public housing residents, while expanding demolition-area redevelopment and replacement construction delivers large gains for the broader city population through general-equilibrium channels.

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1 Introduction

Place-based redevelopment and community investment programs are a consequential form of government intervention for addressing geographic inequality (Moretti, 2024; Neumark and Simpson, 2015; Kline and Moretti, 2014). These programs target specific geographic areas—typically neighborhoods designated as economically distressed—and deliver support through tax incentives, grants, and public investment rather than transfers keyed to individual income. In the United States, estimates suggest that place-based development involves tens of billions of dollars in ongoing state and local incentives (over $80 billion annually) alongside a recent federal surge of more than $70 billion in explicitly place-based industrial policy programs authorized through 2027 (Story, 2012; Muro et al., 2022).

Policymakers have argued that place-based investments can revitalize distressed areas by improving housing, amenities, and economic opportunities, and have often framed redevelopment as a strategy to promote "mixed-income" neighborhoods and expand opportunities for low-income residents (U.S. HUD, 2025; U.S. GAO, 2006). Critics contend, however, that redevelopment frequently displaces incumbent residents and generates broader equilibrium responses—changes in neighborhood composition and prices—that shift the gains of revitalization toward higher-income households (e.g., Smith, 2005; Freeman, 2011; Zuk et al., 2018). Popular media accounts similarly emphasize the costs borne by displaced households, describing cycles of dislocation in which relocation assistance is limited and promised replacement housing is delayed or never delivered (e.g., Schuetz, 2025; Dumke, 2022; Dillon and Poston, 2021). Assessing the welfare effects of these place-based investment programs therefore requires understanding both how households 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. HOPE VI targeted developments that met standards of extreme physical disrepair, economic distress, and social disorganization. Over a nearly two-decade period, the program spent more than $6 billion to transform disadvantaged areas through demolition and redevelopment, while typically supporting the relocation of original residents by providing them with standard rental housing vouchers (U.S. GAO, 2007). Our analysis is organized around three linked questions. First, what are the welfare costs borne by public-housing residents who are forced to relocate? Second, how do demolitions and redevelopment shape neighborhood composition and housing prices? Third, how are the resulting welfare gains and losses distributed across households in the wider city?

Our empirical setting is Chicago, which had one of the nation's largest public housing systems and experienced an intensive wave of HOPE VI-era demolitions. We estimate a structural model of neighborhood choice using two complementary sources of information. Our first main source is administrative records from Illinois that track the location of Chicago's public housing residents before and after demolition, allowing us to quantify the welfare costs of forced displacement and other channels. The second main source is the Census and American Community Survey (ACS) tract-level data for Cook County to both describe citywide sorting patterns and housing-market outcomes as well as to estimate neighborhood demand for residents outside public housing. Finally, we use our estimated model to study the equilibrium effects of demolitions on neighborhood composition, prices, and welfare for displaced and non-displaced residents throughout the city.

We begin our analysis by documenting a set of stylized facts in our Chicago data that speak directly to the impacts of HOPE VI—both for households compelled to relocate and for the broader set of Chicago residents. In the administrative records, we show that public-housing residents who are forced to move because their building was demolished do not typically relocate to highly advantaged neighborhoods, even with voucher-based assistance: moves are often incremental and many destination neighborhoods are economically disadvantaged.[1] Consistent with sizable moving frictions, voluntary mobility is limited in our sample of similar public-housing households who were not displaced: just 20 percent of the control group of households move across neighborhoods in the years following the initial wave of public housing demolitions. The Census and ACS data further show that demolitions represented large local shocks to the housing stock. Between 2000 and 2010—when most demolitions occurred—areas targeted for redevelopment experienced sharp neighborhood change and sizable increases in rents and housing values, alongside evidence of new construction and reinvestment.

These stylized facts motivate a quantitative spatial-equilibrium model in which the welfare impacts of HOPE VI are shaped by two forces: the tradeoff between displacement costs and limited scope for neighborhood upgrading for forced movers, and broader equilibrium adjustments in prices and sorting that extend beyond demolition tracts. In our model, households choose neighborhoods based on a range of factors including the cost of housing, the economic and demographic composition of neighbors, and the presence of public housing itself. Demolition and redevelopment can therefore affect welfare directly—by changing the set of neighborhoods available to displaced households—and indirectly by shifting neighborhood desirability, altering equilibrium prices, and inducing re-sorting that changes endogenous neighborhood characteristics. Because HOPE VI compels some households to move, welfare depends not only on origin and destination attributes but also on the costs of displacement. Our framework formalizes these channels in a neighborhood choice model with an elastic housing supply, allowing prices, neighborhood composition, and amenities to adjust endogenously and translating observed relocation decisions and market-wide changes in sorting into welfare impacts. We build on structural neighborhood-choice models that discipline moving costs and neighborhood demand, adapting them to a setting with compelled relocation (e.g., Bayer, Ferreira and McMillan, 2007; Galiani, Murphy and Pantano, 2015; Bayer et al., 2016).

For identification, our empirical strategy relies on two distinct sources of variation. First, we use the individual-level administrative records for public housing residents and exploit quasi-experimental variation from the program's initial wave of demolitions: some buildings were

[1]The evidence is consistent with findings in prior work on HOPE VI and Chicago demolitions (Jacob, 2004; Chyn, 2018).

torn down while nearby ones within the same project remained standing for several years. This timing generates a treated group that was displaced by demolition and a comparable control group that was not. The control group is essential because their baseline mobility pins down voluntary move rates and thus identifies moving-cost parameters. Moreover, since the exposure to demolition is quasi-random, treated and control households face rent differences that are plausibly exogenous. Conditional on moving from similar origin tracts, differences in destination choices between the two groups reveal housing price sensitivity through the contrast between voucher-subsidized and market rents. Second, to identify the remaining parameters in our model, we estimate preferences for endogenous neighborhood demographics using shift-share instrumental variables that combine information on the pre-existing distribution of particular types of residents from particular birth states/countries and nationwide shifts in population across birth locations. For the broader set of residents, we use Census and ACS data along with these same instrumental variables to estimate preferences, while calibrating moving costs and the housing supply elasticity using external benchmarks to close the model.

We begin our preview of the results by summarizing estimated neighborhood demand and relocation frictions—key inputs into the welfare analysis. For public-housing residents (essentially all of whom were Black), the estimates indicate high sensitivity to housing costs, a distaste for neighborhoods with public housing (consistent with accounts of the physical disrepair of buildings and social distress in these areas), a preference for living in neighborhoods with same-race and higher-income residents, and substantial utility costs of moving. For example, we estimate that public-housing residents would be willing to pay \$148 more in annual rents for a 1 percentage point decrease in the share of the housing stock that is public housing. For the broader population, we estimate preferences that are qualitatively similar.

Our main welfare results follow from combining these estimates with housing supply responses and solving for equilibrium outcomes with and without HOPE VI demolitions. For displaced public-housing residents, we estimate large welfare losses implying that any benefits from neighborhood upgrading and voucher-enabled choice are insufficient to fully compensate displacement costs. For the broader population, the equilibrium response reallocates gains and losses across groups as prices and neighborhood composition adjust: renters are made worse off because citywide increases in housing prices due to demolitions outstrip the value of the changing amenities, while homeowners see gains because higher prices translate into increased income. Combined, these forces generate welfare losses for lower-income households—especially racial minorities—and welfare gains for higher-income residents, reinforcing economic inequality in the impact of demolitions. In the aggregate, we estimate an overall loss from demolitions worth about −\$5 million in rent equivalent units, largely driven by the impacts on displaced public housing households.

To clarify the mechanisms behind these patterns, we conduct parallel decompositions for displaced households and for the broader population. For public-housing residents, the decomposition highlights two distinctive features of the setting. First, when abstracting from forced relocation, public housing demolitions increase the welfare of public-housing residents, who are especially sensitive to the disamenities of these buildings. Second, demolitions both compel households to relocate and expand their neighborhood options via vouchers. While vouchers generate meaningful gains, they are not sufficient to offset the large utility costs of displacement. For the broader population, we decompose effects into the direct value of neighborhood change holding prices and composition fixed and the indirect general-equilibrium response operating through re-sorting, endogenous amenities, and price adjustments, which helps explain why renters often lose while homeowners benefit from appreciation.

A natural next question is how alternative housing policies can shape these welfare consequences. To address this, we use the estimated model to conduct counterfactual simulations that shift policy along multiple margins. On the demand side, we study two targeted interventions primarily aimed at displaced public-housing residents: increasing the generosity of voucher assistance by relaxing the payment standard tied to local rent benchmarks and providing intensive housing counseling to reduce the costs of relocation. In terms of the supply side, we vary the scale of new construction in demolition-targeted neighborhoods. Raising voucher generosity by 10 or 20 percent increases the share of neighborhoods in which the voucher can fully cover market rents and mitigates—but does not eliminate—welfare losses for displaced households. Reducing moving costs through intensive counseling produces much larger attenuation of the negative impacts of demolition. Because displaced households account for a substantial share of the total welfare effect, these targeted demand-side interventions can materially shift the overall welfare impact of demolitions. Our housing redevelopment intervention generates the largest aggregate welfare gains—concentrated almost exclusively among the broader set of Cook County residents—by directly changing local prices and neighborhood composition through general-equilibrium adjustment.

Overall, the main contribution of this paper is to provide the first welfare and incidence evaluation of a large place-based redevelopment shock that explicitly incorporates forced relocation within a quantitative spatial-equilibrium framework. A growing literature uses spatial equilibrium and neighborhood-sorting models to translate place-based interventions and other large localized shocks—such as infrastructure upgrades, transit investments, or slum upgrading—into welfare changes that incorporate endogenous re-sorting and price responses (e.g., Harari and Wong, 2018; Tsivanidis, 2019; Balboni et al., 2020; Redding and Sturm, 2024). We complement this important work by bringing to the forefront a margin that is often difficult to observe directly in place-based settings: the welfare costs borne by households who are compelled to move. By combining administrative micro data that track displaced public-housing residents with citywide Census/ACS outcomes, we can estimate moving costs and neighborhood demand for the displaced group and then embed these households in a citywide equilibrium in which prices, composition, and amenities adjust endogenously. This structure allows us to quantify how displacement losses interact with market-wide adjustments to determine the distribution of gains and losses across displaced households and the broader set of residents.

We also contribute to studies of HOPE VI—one of the largest urban redevelopment programs in U.S. history—by connecting evidence on relocation outcomes to a welfare accounting that makes displacement costs explicit. Prior work on Chicago demolitions shows that many households relocate using vouchers and that children's outcomes improved in the long-run (Jacob, 2004; Chyn, 2018; Haltiwanger et al., 2024). Policy evaluations likewise document relocation patterns and subsequent mobility among HOPE VI relocatees (e.g., Popkin et al., 2002). At the same time, qualitative and mixed-methods studies emphasize housing uncertainty and the loss of place-based social ties during relocation (e.g., Bennett, Smith and Wright, 2006; Manzo, Kleit and Couch, 2008), echoing broader evidence that moves from disadvantaged neighborhoods can impose substantial non-monetary costs even when housing conditions improve (e.g., Barnhardt, Field and Pande, 2017). We incorporate these mechanisms directly: we estimate neighborhood demand and moving costs for public-housing households using quasi-experimental variation, decompose welfare into relocation frictions versus voucher-enabled choice gains, and then trace how equilibrium price and sorting responses shape welfare for the broader market.

2 Background: Public Housing and HOPE VI in Chicago

2.1 The Public Housing System in Chicago Before HOPE VI

At the beginning of the 1990s, Chicago had the third largest public housing system in the United States. The Chicago Housing Authority (CHA), a local government agency overseen by a mayor-appointed board, owned and managed the city's subsidized housing stock, providing assistance to more than 50,000 residents (CHA, 2000). The CHA's portfolio included high-rise housing developments (known as "projects") and smaller-scale scattered-site residential buildings. High-rise projects typically comprised multiple apartment buildings built in close proximity, with many buildings containing approximately 75 to 150 housing units. The largest high-rise developments, such as the Robert Taylor Homes and Cabrini-Green, housed tens of thousands of residents and became national symbols of large-scale public housing (Hunt, 2001; Austen, 2018).

Public housing in Chicago was concentrated in neighborhoods on the South and West sides of the city, largely in predominantly Black neighborhoods, reflecting postwar siting decisions that reinforced long-standing residential segregation. In the 1950s and 1960s, much of the high-rise stock was built through slum clearance and urban renewal initiatives, and proposed sites in White neighborhoods often faced intense political resistance—opposition that helped concentrate family public housing in already-segregated areas (Hunt, 2009; Hirsch, 2009). As a result, high-rise housing was clustered in a relatively small set of neighborhoods, narrowing the set of locations available to households who entered project-based assistance.

Within this spatial context—and under federal eligibility and rent rules—CHA public housing primarily served an overwhelmingly Black, extremely low-income resident population. Eligibility was limited to households with incomes at or below 50 percent of the local area median income, and tenant rent contributions were income-based—typically about 30 percent of adjusted income—so required payments scaled with ability to pay. In practice, residents during this period were far poorer than the income-limit ceiling: average household income was about $7,000 (well below Chicago's 1990 median family income of $44,000) (U.S. Department of Housing and Urban Development, 1990; Popkin et al., 2000). The resident population was also predominantly single-parent and female-headed, with just 6 percent of CHA public-housing households headed by a married couple.

By the end of the 1980s, much of the public housing stock in Chicago was in a state of serious disrepair and required major renovation (Hunt, 2009). This deteriorated state stemmed from both the age of the buildings and chronic underfunding and management failures, which contributed to large maintenance backlogs and major system deficiencies (e.g., lack of reliable heat, plumbing failures, and leaking roofs). 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 public housing units nationwide were "severely distressed," a designation tied to physical deterioration. The Commission noted that many developments had reached a point where they no longer provided safe, sanitary, and decent housing, and it defined severe distress in operational terms that included extensive deferred maintenance (e.g., work-order backlogs) and major building-system failures (U.S. National Commission on Severely Distressed Public Housing, 1992).

2.2 *Demolitions and Redevelopment*

Chicago's large-scale demolition of high-rise public housing emerged as a reaction to the severe physical distress of the housing stock and was facilitated by a policy environment that increasingly treated redevelopment as the only viable path forward (Popkin et al., 2000; Hunt, 2009). Federal policy played a central enabling role through HUD's HOPE VI program, launched in 1992 to support the revitalization of severely distressed public housing nationwide by providing hundreds of competitive grants to local housing authorities (Congressional Research Service, 2011). HOPE VI grants could be used not only for demolition and new construction, but also for resident relocation and supportive services, explicitly linking physical redevelopment to the transition of incumbent households.

Chicago was one of the largest of the program's recipients, and HOPE VI became a key funding source for the city's redevelopment agenda. Between 1996 and 2003, the city received $83.4 million in HOPE VI grants designated for building demolition-related activities (Aliprantis and Hartley, 2015). Building on demolition activity already underway in the mid-1990s, the CHA later formalized its agenda through the Plan for Transformation, launched in 1999 and authorized through a CHA-HUD agreement in early 2000 (Chicago Housing Authority, 1999). The Plan laid out a blueprint to remove much of the high-rise family housing stock and replace it with mixed income redevelopment combining replacement public-housing units with affordable and market-rate housing.

Demolition and the relocation of public housing residents proceeded gradually, with roughly

80 percent of demolitions occurring between 2000 and 2010 (see Appendix Figure A.1). Residents displaced by demolition typically received Section 8 housing voucher offers, often alongside administrative support and relocation assistance (Vale and Graves, 2010). While vouchers expanded the set of neighborhoods and units that families could access relative to project-based assistance, they did not eliminate affordability constraints. In particular, the voucher subsidy was capped by a payment standard tied to local Fair Market Rents (FMRs), so leasing a unit with rent above the payment standard required paying the difference out of pocket.[2] Importantly, vouchers preserved the same basic income-based rent formula as public housing—assisted households typically paid about 30 percent of adjusted income toward rent and utilities—so the transition from public housing to vouchers did not mechanically raise disposable income or required rent burdens.

In practice, the voucher-based transition was often difficult to implement and attracted scrutiny from the outset, even among households that desired access to improved housing quality and safer environments. Contemporary interviews and focus groups conducted during the Plan for Transformation indicate that many voucher-assisted households remained anchored to their origin neighborhoods for schools, childcare, churches, and services, and many expressed a desire to return locally even when they were uncertain about returning to CHA housing (Williams, Fischer and Russ, 2003). More broadly, HOPE VI-era evaluations and qualitative studies—often based on modest samples—emphasize that the relocation process involved substantial search and administrative frictions and that feasible options could be constrained by local market conditions and by program rent ceilings, since voucher assistance is capped by payment standards tied to Fair Market Rents (Popkin, 2007; Manzo, Kleit and Couch, 2008; U.S. Department of Housing and Urban Development, 2025).

Finally, neighborhood redevelopment under the Plan for Transformation proved uneven as demolition proceeded faster than new construction in many neighborhoods. Popular media accounts describe long stretches of cleared land remaining underutilized for years and rebuilding efforts falling notably short of initially planned targets (Dumke, 2017; Bittle, Kapur and Mithani, 2017). Consistent with these accounts, summary statistics from land-use data indicate that redevelopment lagged for a nontrivial share of former project sites: Appendix Table A.1 shows that 38 percent of parcels where public housing was demolished remained vacant and undeveloped in 2010.[3] This slow pace of redevelopment reduced the scope for displaced households to re-establish themselves in or near their former communities and shaped how the demolition shock translated into longer-run neighborhood change.

[2]The voucher subsidy equals the difference between the program's maximum allowable rent (the payment standard, based on local FMRs and bedroom size) and the household's required rent contribution. During our period, FMRs were based on roughly the fortieth percentile of rents in the local private-market distribution.

[3]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. These land use statistics indicate that land redeveloped by 2010 primarily contained residential housing (40 percent), although some was also occupied by businesses (8 percent) and institutions (4 percent, mostly schools and government buildings). Even by 2015, there was minimal additional progress as the share of vacant land stood at 35 percent.

3 Data and Analytic Samples

Our analysis combines multiple data sources to create two main analytic samples. First, we link Chicago Housing Authority (CHA) building records to Illinois Department of Human Services (IDHS) social assistance case files, allowing us to identify public-housing households before public housing demolition and track their locations over time. Second, we use tract-level tabulations from the decennial Census and the ACS, together with CHA demolition records, to measure neighborhood composition and housing-market outcomes throughout Cook County and to quantify how demolitions reshaped neighborhood prices and sorting in spatial equilibrium. The remainder of this section details the construction of each sample and the key variables we use.

Our analysis of public housing households focuses on the subset of CHA buildings and residents exposed to the initial HOPE VI-era demolition wave (1995–1998). More specifically, we use CHA building records to create a list of family (non-senior) public housing high-rise buildings, defined as having 75 units or more. In general, low- and mid-rise buildings did not experience the same type of abrupt demolition as high-rises. For each building, we identify dates of building closure due to public housing demolition using administrative records on occupancy, following the approach in Jacob (2004) and Chyn (2018). We define a comparison group of buildings by identifying those that do not close during the initial wave of public housing demolition. For example, in the Robert Taylor Homes project, Building 1 was slated for demolition in 1995 while other high-rises in Robert Taylor were left untouched at the time and remained open for many years (on average eight years) after the initial building demolitions. We then match the set of demolished and comparison group buildings by address to the IDHS case files to identify resident adults living in public housing in the year prior to building closure. This yields an analytic sample of adults observed in the IDHS records whom we can follow over time.[4] For each individual, we use address records over the three years following building closure (for individuals in demolished buildings) or the first nearby building closure in the same project (for individuals who are not displaced) to measure whether they remain at their baseline public-housing address in subsequent periods and, if they exit, their first observed address outside public housing. Because comparison buildings remain open during the initial wave, we observe both continued residence and voluntary exits for the comparison group, while treated residents necessarily exit at demolition. Our main analytic sample comprises 2,966 adults for whom we observe both the baseline address and the first post-demolition address. Appendix Table A.2 provides baseline summary statistics.

The next main sample focuses on neighborhoods throughout Cook County using tabulations from the decennial Census and the ACS for 1990, 2000, 2010, and 2016. Neighborhoods are defined as census tracts harmonized to 2010 boundaries using crosswalks from the Longitudinal Tract Data Base (Logan, Xu and Stults, 2014). We link these tract-level data to CHA demolition activity using building records compiled by Sandler (2017).[5] The Census/ACS tabulations measure neighbor-

[4] Chyn (2018) shows no detectable impacts of public-housing demolitions on social-assistance participation, mitigating concerns that tracking households through IDHS mechanically conditions on an outcome affected by demolition.

[5] The building closure timing used in our individual-level sample is based on CHA occupancy records that end in

hood demographics (race/ethnicity and income) and housing-market outcomes, including median rents, median home values, and housing stock characteristics.[6] Our tract-level sample includes all tracts in Cook County to allow for spillovers beyond Chicago city limits; after dropping tracts with missing key variables, the analysis includes 1,240 tracts.[7] On average, each tract contains roughly 4,000 residents.

4 Stylized Facts: Public Housing Displacement and Neighborhood Change

This section documents a set of descriptive facts that motivate our equilibrium analysis of HOPE VI in Chicago. We begin by characterizing the demolition shock itself: demolitions were spatially concentrated and large relative to the pre-existing housing stock in a subset of neighborhoods. We then describe what happened to public-housing households exposed to the initial demolition wave. Demolition-induced displacement generated sharply elevated cross-neighborhood mobility, and displaced households moved to less disadvantaged neighborhoods on average, but remained far from the typical neighborhood in Cook County. Finally, we show that neighborhoods experiencing more intense demolition underwent substantial changes in demographics and housing-market conditions—changes that grow over the longer run and therefore are unlikely to reflect only the mechanical relocation of former public-housing residents. Together, these patterns imply that our framework should contain two key features: (i) moving frictions that shape relocation outcomes and household welfare alongside (ii) endogenous neighborhood composition and rents adjustments in response to large local shocks to the housing stock.

4.1 The Geography of Public Housing Demolitions

We start by summarizing where demolitions occurred and how large the resulting shocks were at the neighborhood level. Public-housing demolition activity was not diffuse across the city: it was concentrated in a relatively small set of 59 census tracts (5 percent), predominantly on Chicago's South and West sides, reflecting the pre-HOPE VI geography of high-rise public housing described in Section 2. Figure 1 maps cumulative demolished public-housing units by census tract, highlighting substantial spatial heterogeneity in demolition exposure across neighborhoods.

Because the consequences of demolition for neighborhood outcomes depend on the scale of demolished units relative to the local housing market, we also summarize variation in demolition intensity normalized by the pre-existing housing stock. Panel B plots demolitions as a share of 1990 occupied housing units for tracts that experienced demolition. The distribution is right-skewed: while many tracts experienced demolition activity that was modest relative to the 1990 stock, a sizable subset of neighborhoods experienced demolitions that represented a large fraction

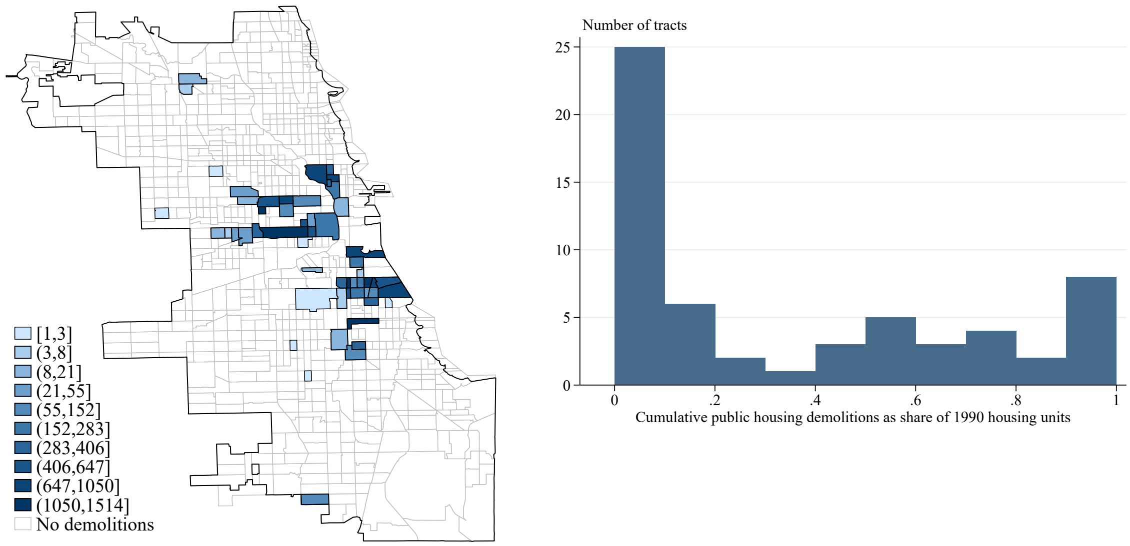
2000, which limits their usefulness for measuring demolition timing over the full Plan for Transformation period. The Sandler (2017) CHA records report the number of public-housing units in each building and the structural demolition date. We geocode building addresses, map them to census tracts, and construct tract-level measures of demolition intensity based on the number of units demolished.

[6] We use 5-year ACS tabulations for 2008–2012 (labeled 2010) and 2014–2018 (labeled 2016); 2010 population counts come from the decennial Census. Throughout, we use household counts and the race/ethnicity of the household head.

[7] We drop two tracts with public housing due to missing outcomes.

Figure 1: Spatial Variation in Public Housing Demolitions in Chicago

(a) Cumulative Demolitions by Tract (b) Distribution of Demolitions as Share of Housing



Notes: Panel A displays the cumulative number of public housing units that were demolished between 1995 and 2010 in each census tract in Chicago. 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.

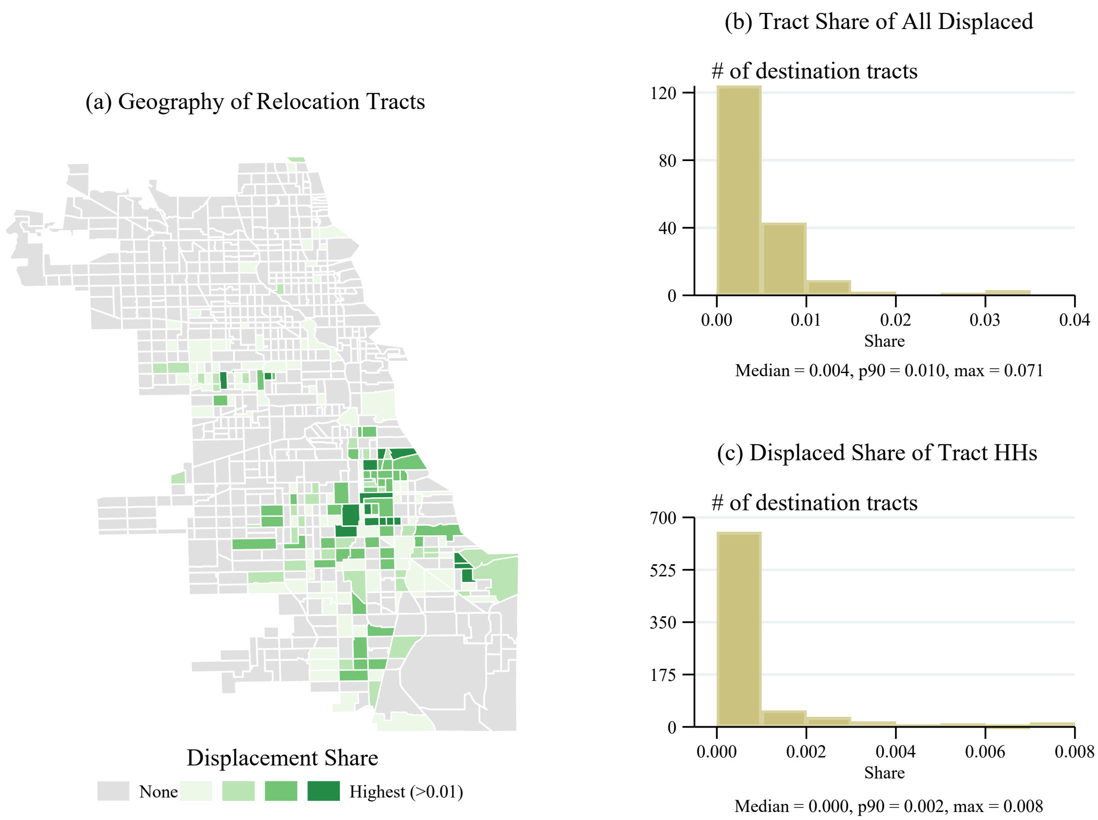
of local occupied housing units. These patterns underscore that HOPE VI operated not only as a household-level displacement event, but also as a large local shock to neighborhood housing supply and the structure of the housing stock.

4.2 Public Housing Displacement and Relocation Outcomes

Next, we turn to our main analysis of public housing households exposed to the initial demolition wave (1995–1998). Our linked administrative data allow us to measure whether baseline residents remain at their public-housing address after the closure of their building and, if they exit, the location of their first observed address outside public housing (Section 3). We use this information to summarize relocation outcomes for displaced households, benchmarking them both against residents in comparison buildings that remained open during the same period and against the broader set of households in Cook County.

We begin by describing where displaced public-housing households relocate in Figure 2. Panel A maps the share of displaced residents relocating to each Chicago census tract, and Panel B shows how those destination shares are distributed across tracts. The pattern reflects a wide set of realized destinations alongside pronounced concentration in broad areas of the city. Displaced households are observed relocating to 191 distinct tracts (out of 788 tracts citywide and 1240 tracts countywide). Relocation shares are far from evenly distributed and are heavily concentrated in a limited subset of neighborhoods. The histogram underscores this skewness: among destination tracts, the

Figure 2: Relocation Tracts for Displaced Public-Housing Households



Notes: Panel A maps, for each Chicago census tract, the share of displaced public-housing households in our administrative records whose destination tract is that tract (i.e., the number of displaced households relocating to tract *j* divided by the total number of displaced households with an observed destination in Chicago). Tracts that receive no displaced households are shaded in gray. Darker shading indicates larger destination shares; the highest category corresponds to tracts receiving more than 0.01 (one percent) of displaced households, as indicated in the legend. Panel B shows the distribution of tract-level destination shares across tracts. Panel C shows the distribution of displaced arrivals scaled by tract size, defined as the number of displaced households relocating to tract *j* divided by the total number of households in tract *j* in 2000; this panel includes all Chicago tracts, so the mass at (near) zero reflects tracts that receive no displaced households.
Source: Authors' calculations using Chicago Housing Authority administrative records matched to census tracts, and tract household counts from the Census/ACS.

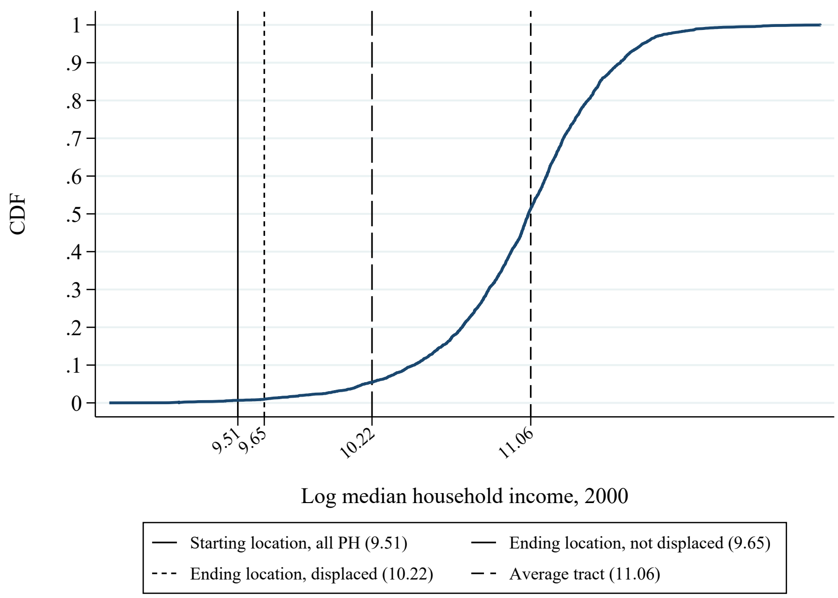
median tract receives 0.004 of displaced households (0.4 percent), the 90th percentile receives 0.01 (1 percent), and the maximum tract receives 0.071 (7.1 percent), indicating that most tracts receive very small shares while a small number receive comparatively large inflows. Spatially, the tracts with the largest destination shares are concentrated on the South and West sides—often in areas near the originating public-housing communities—with comparatively little relocation to many other parts of the city. Consistent with this geographic clustering, the median relocation distance in Cook County is 2.5 miles (p90: 8.1 miles), and 75 percent of households relocate within five miles of their baseline tract. This destination structure is informative for our formal neighborhood choice analysis: relocations span many distinct tracts—providing the variation needed to learn how households trade off neighborhood attributes—yet remain highly concentrated and spatially local, consistent with constraints such as proximity to prior neighborhoods playing an important role alongside neighborhood characteristics.

To place these destination patterns in a broader context, Figure 3 plots the cumulative distribution function (CDF) of tract-level log median household income in Cook County in 2000 (weighted by the number of households). The vertical lines overlay reference values for the average income of (i) households' baseline tracts in public housing (measured prior to building closure), (ii) the post-demolition tracts of residents in comparison buildings, (iii) the destination tracts of displaced households, and (iv) the average tract in Cook County. Displaced households relocate to higher-income neighborhoods on average relative to their baseline public-housing tracts, indicating some degree of neighborhood upgrading after leaving project-based assistance. At the same time, their typical destination neighborhood remains far below the Cook County distribution, highlighting that relocation outcomes are concentrated in a limited segment of the countywide neighborhood distribution. In other words, displacement shifts households rightward in the neighborhood-income distribution, but the shift is limited relative to the full range of neighborhoods in the county.

Figure 3 also highlights how the relocation environment facing displaced households differs from that of residents in comparison buildings. In particular, displaced households' destination neighborhoods are higher-income on average than the post-period neighborhoods of residents in comparison buildings. This gap is consistent with the idea that the demolition relocation process—including voucher offers and administrative support—can facilitate larger neighborhood upgrades than typically occur absent forced displacement, even as both groups' locations remain well below the countywide distribution.

Table 1 complements Figure 3 by summarizing destination neighborhoods along multiple dimensions. Column 2 confirms the substantial baseline disadvantage of public-housing neighborhoods emphasized above: relative to the Cook County average in column 1, baseline tracts have much lower median income and rent and are overwhelmingly Black. The table also highlights that changes in neighborhood racial composition are comparatively modest relative to the economic upgrading associated with displacement. The average Black share falls from 0.99 in baseline tracts to 0.91 in displaced households' destination tracts, a decline of 8 percentage points (and 8 percent

Figure 3: Starting and Ending Tracts in the Cook County Income Distribution



Notes: This figure is a plot of the CDF of tract-level log median household income in Cook County in 2000, weighted by the number of households. The vertical lines correspond to the tract-level average of log median household income for the public-housing sample in their starting location, the average for households who were not displaced by demolitions in their ending location, the average for households who were displaced by demolitions in their ending location, and the average for all households in Cook County.

relative to the baseline tracts), leaving destination neighborhoods still overwhelmingly Black and far above the Cook County average. By contrast, average log median household income rises from 9.51 in baseline tracts to 10.22 in destination tracts—a 0.71 log-point increase, corresponding to roughly a 103 percent increase in neighborhood median income—and log median rent rises from 5.63 to 6.41 (about 118 percent). The relative stability of neighborhood racial composition, even alongside sizable income upgrading, suggests that neighborhood composition is a salient dimension of the locations households select into after displacement. The last row highlights that 38 percent of displaced households relocate to tracts that still contain public housing, reinforcing that even when households exit project-based assistance, relocation outcomes remain closely tied to a limited set of neighborhoods.

Table 1: Neighborhood Characteristics, All Households and Public Housing Sample

Neighborhood characteristic	All Cook County households (1)	Public housing households: Starting location, all PH households (2)	Public housing households: Ending location, not displaced by demolitions (3)	Public housing households: Ending location, displaced by demolitions (4)
Share Black	0.26	0.99	0.97	0.91
Share Hispanic	0.14	0.01	0.01	0.04
Log median household income	11.06	9.51	9.65	10.22
Log median rent	6.90	5.63	5.76	6.41
Had public housing units	0.05	1.00	0.87	0.38

Notes: Column 1 reports average neighborhood characteristics, weighted by the number of all households that live in each tract in Cook County in year 2000. Column 2 reports average characteristics for our public housing household sample in their starting location (when all are observed in public housing). Columns 3 and 4 report average characteristics approximately three years later, for the group who was not displaced by demolitions during this period (column 3) and the group who was (column 4).
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

We close by summarizing how sharply demolition altered residential mobility. Appendix Figure A.2 reports cross-tract migration rates over the post-demolition window. The figure shows that displacement produces a sharp increase in cross-neighborhood moves: the vast majority of displaced households move across census tracts, while mobility is much lower for residents in comparison buildings over the same period. This low mobility in the comparison group helps explain why their average location shifts little in Figure 3: over short horizons, most residents whose buildings remain open do not change census tracts. The contrast is informative more broadly because it underscores that the demolition wave generated an unusually large out-migration flow, rather than simply accelerating moves that would have occurred anyway. The final bar in Appendix Figure A.2 provides an additional benchmark from a broader disadvantaged population, reinforcing that demolition-induced moves are elevated relative to typical mobility rates in high-poverty contexts.

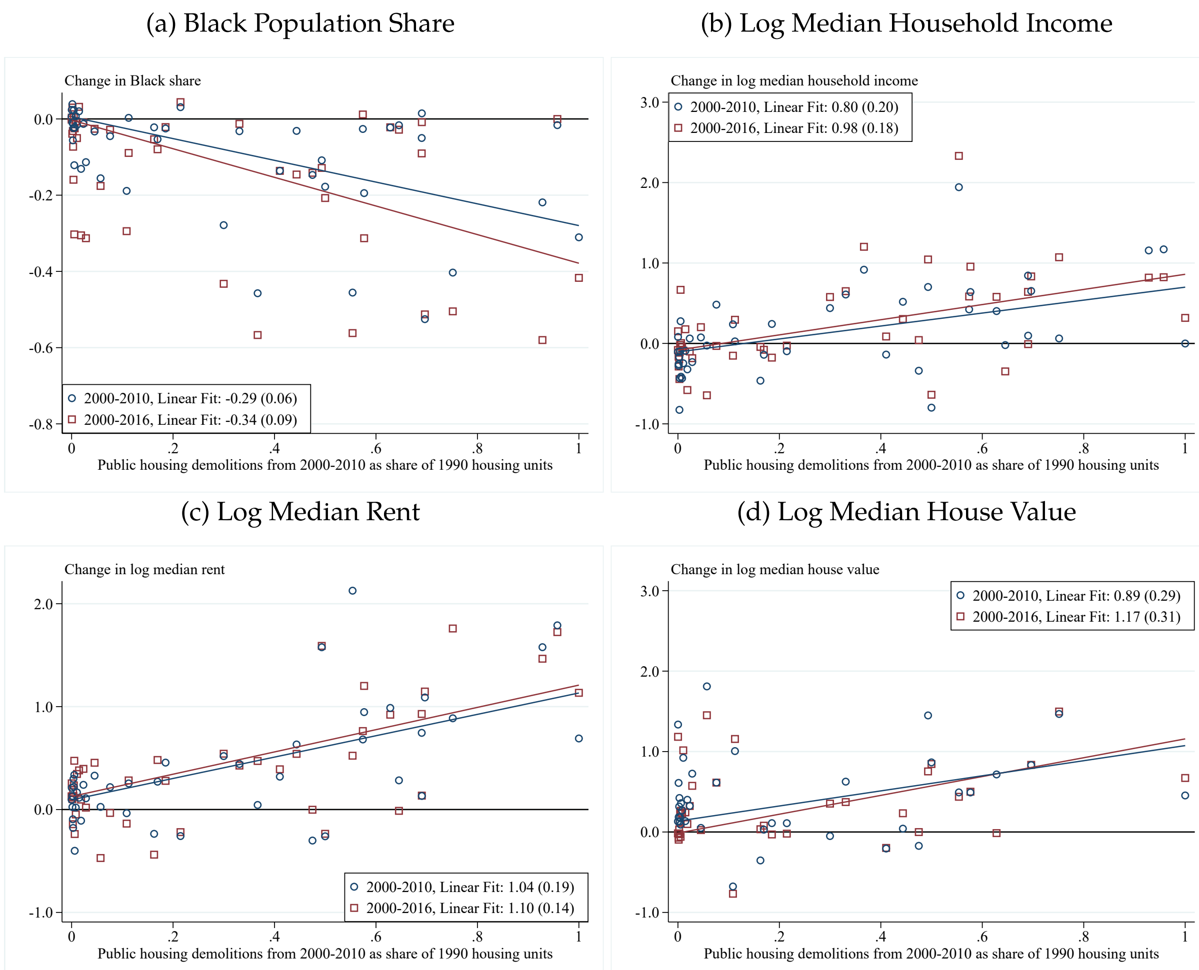
4.3 *Neighborhoods and Housing Markets After Demolitions*

We next document how neighborhood composition and housing-market conditions evolved in areas exposed to HOPE VI-era demolitions. Because demolitions were spatially concentrated and large relative to the pre-existing housing stock in a subset of tracts, they represent a sizable local shock to local housing supply and neighborhood amenities. The demolition of public-housing buildings was a major undertaking and produced temporary disamenities associated with this construction effort. After these buildings were taken down, the land was redeveloped in about two-thirds of cases and left vacant in the remaining ones (Appendix Table A.1). Moreover, the departure of public-housing residents mechanically changes the demographic and socioeconomic composition of affected neighborhoods in the short run, while demolition and subsequent redevelopment can also reprice neighborhoods and induce broader resorting among households beyond those originally living in public housing. To describe how neighborhoods changed following this shock, we relate changes in tract characteristics to the cumulative intensity of public-housing demolition. Specifically, we present binned scatter plots of changes in census-tract outcomes between 2000 and 2010 against cumulative demolitions from 2000–2010, measured as the number of public-housing units demolished in the tract divided by the tract's occupied housing units in 1990. We also report analogous changes between 2000 and 2016 to capture longer-run patterns.

Figure 4 summarizes how neighborhood demographics and housing markets change with demolition intensity. Panels A and B show that higher-demolition tracts experience larger declines in the Black population share and larger increases in median household income. The displacement of public-housing residents—an overwhelmingly Black and very low-income population—mechanically accounts for part of these short-run shifts in tract composition. The longer-run results, however, suggest that neighborhood transition extends beyond this immediate compositional change. In particular, the relationships are noticeably stronger when measured from 2000 to 2016 than from 2000 to 2010, consistent with additional resorting and neighborhood change unfolding over time. Panels C and D show corresponding changes in local housing markets: rents and housing values grow faster in higher-demolition tracts, indicating substantial repricing in areas more exposed to demolition. Housing values are especially informative in this setting because they embed forward-looking expectations about neighborhood change and are not mechanically affected by the removal of low-rent public-housing units in the same way that median rents can be. Appendix Figure A.3 reports complementary results for other demographic groups and for changes in the age distribution of the housing stock. The patterns reinforce the evidence of neighborhood transition and repricing shown in Figure 4, and they are consistent with substantial housing-stock turnover in higher-demolition areas.

Taken together, these neighborhood-level patterns motivate the equilibrium framework we develop in the sections that follow. They indicate that HOPE VI operated not only through direct displacement but also through broader adjustments in neighborhood composition and prices that are likely to affect many households beyond those living in public housing. In our model, we therefore allow households to sort across neighborhoods with heterogeneous preferences, while

Figure 4: Neighborhood Change and Public Housing Demolitions, 2000–2010 and 2000–2016



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 for all tracts in Cook County. Each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition, with a single bin for tracts with no demolitions. We winsorize the public housing demolition share variable from above at 1 for 3 tracts.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

neighborhood composition and prices jointly respond to changes in housing supply and local conditions induced by demolition and redevelopment.

5 A Model of Sorting Across Neighborhoods

The demolition of public housing could affect households, neighborhoods, and housing markets in several ways. Households who are forced to relocate experience costs associated with this displacement and face lower out-of-pocket rents through their receipt of a Section 8 voucher subsidy. Households throughout Chicago might also make different location choices due to changes in the presence of public housing or in socioeconomic and demographic characteristics of a neighborhood's residents following demolitions. These shifts in residential sorting can, in turn, change equilibrium housing prices and reshape endogenous neighborhood characteristics that are jointly determined with sorting. 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; Galiani, Murphy and Pantano, 2015) that features endogenous amenities with a model of housing supply.

5.1 Housing Demand

We begin by specifying a flexible model for household location choice as follows. Households choose among J neighborhoods within Cook County, indexed by $j \in \{1, \ldots, J\}$, and an outside option $j = 0$ corresponding to living outside Cook County (within Illinois). 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.

At the start of period t, each household i is endowed with an origin location, denoted by $\ell \equiv j(i, t-1)$, that summarizes its location in a baseline period. We classify households into K-many types based on demographic and socioeconomic characteristics, including whether a household resides in public housing at $t-1$ (our "public housing" type). Household i of type k decides whether to move to a new neighborhood by solving the following problem:

$$\max_j V^k_{ijt|\ell} = \delta^k_{jt} + \alpha^k_p \ln(p_{ijt}) - MC^k_{\ell,j} + \epsilon^k_{ijt}, \tag{1}$$

where δ^k_{jt} is the component of indirect utility for neighborhood j that is common to all households of group k, p_{ijt} is the rental price of housing (which might vary across households due to Section 8 housing vouchers), $MC_{\ell,j}$ is the cost of moving from ℓ to j, and ϵ^k_{ijt} is an idiosyncratic shock that is assumed to be an i.i.d. type I Extreme Value. The common component of indirect utility is:

$$\delta^k_{jt} \equiv \alpha^k_b b_{jt} + \alpha^k_h h_{jt} + \alpha^k_{Inc} \ln(Inc_{jt}) + \alpha^k_{PH} PH_{jt} + \theta^k x_{jt} + \xi^k_{jt} \quad \forall\, j \geq 1, \tag{2}$$

where 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, $\alpha^k \equiv (\alpha_p^k, \alpha_b^k, \alpha_h^k, \alpha_{Inc}^k, \alpha_{PH}^k, \alpha_x^k)$, moving costs, $MC_{\ell,j}^k$, as well as neighborhood unobserved quality, ξ_{jt}^k, may differ arbitrarily across groups. This heterogeneity allows different populations to value neighborhood attributes and prices differently, which is central for matching observed sorting patterns and for capturing how demolitions can redistribute welfare across groups. In what follows, 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.[8]

Given the distributional assumption on ϵ_{ijt}^k, the probability that a group-k household with origin location $j(i, t-1)$ chooses to live in neighborhood j at time t is:

$$\mathcal{P}_{jt|\ell}^k(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_\mathbf{t}^\mathbf{k}; \alpha^k, MC^k) = \frac{\exp\left(\delta_{jt}^k + \alpha_p^k \ln(p_{ijt}) - MC_{\ell,j}^k\right)}{\sum_{j'} \exp\left(\delta_{j't}^k + \alpha_p^k \ln(p_{ij't}) - MC_{\ell,j'}^k\right)}, \tag{3}$$

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 $k = 1, \ldots, K$:

$$\mathcal{D}_{jt}(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC) = \sum_{k=1}^{K} N_{jt}^k = \sum_{k=1}^{K} N_t^k \sum_{\ell=0}^{J} \mathcal{P}_{jt|\ell}^k(\mathbf{p_t}, \mathbf{b_t}, \mathbf{h_t}, \mathbf{x}_t, \xi_\mathbf{t}^\mathbf{k}; \alpha^k, MC^k) s_{\ell t-1}^k, \tag{4}$$

where N_t^k is the total number of group k households in Illinois, which we take as exogenous, ℓ indexes household locations at time $t-1$, and $s_{\ell t-1}^k$ is the share of group k households that live in neighborhood ℓ at time $t-1$.

There are several important considerations when interpreting our model. First, we view the parameters in equation (2) as reduced-form valuations of neighborhood attributes that may partly reflect 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 or because demographic composition proxies for public goods or consumption amenities that are associated with demographic composition (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 framework allows for several equilibrium responses through rents and demographic composition while keeping tractability. Related, households might also prefer to live in richer neighborhoods because they expect these areas to benefit their children or improve their

[8] When calculating welfare in counterfactual scenarios, we distinguish between renters and homeowners. Changes in the price of housing affect the flow cost that renters and homeowners must pay in each period, while homeowners also are affected through changes in the value of their housing asset implied by the present discounted value relationship.

own employment prospects.[9] Our model does not separately specify or estimate these outcome channels; instead, such consequences affect estimated willingness-to-pay only to the extent they are perceived by households and reflected in observed location choices. 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). Third, we abstract away from uncertainty about the future using the myopic formulation in equation (1), as is done in other work (e.g., Galiani, Murphy and Pantano, 2015; Bryan and Morten, 2019). This allows us to include moving costs and is motivated by the lack of longer panel data in our setting.[10]

5.2 *Housing Supply and Equilibrium with Endogenous Sorting*

We close the model by assuming a housing supply curve. The number of housing units supplied in neighborhood j is an isoelastic function of the price:

$$\mathcal{S}_{jt}(p_{jt}) = \kappa_{jt} p_{jt}^{\psi}, \tag{5}$$

where κ_{jt} is a supply shifter and ψ is the supply elasticity. This functional form is commonly used (e.g., Hsieh and Moretti, 2019; Gaubert et al., 2025) and maps to empirical estimates of housing supply elasticities (Baum-Snow and Han, 2024).

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 $j = 1, \ldots, J$:

$$\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC) = \mathcal{S}_{jt}(p_{jt}^*), \tag{6}$$

$$\frac{\mathcal{D}_{jt}^B(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC)}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC)} = b_{jt}^*, \tag{7}$$

$$\frac{\mathcal{D}_{jt}^H(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC)}{\mathcal{D}_{jt}(\mathbf{p_t^*}, \mathbf{b_t^*}, \mathbf{h_t^*}, \mathbf{x_t}, \xi_\mathbf{t}; \alpha, MC)} = h_{jt}^*, \tag{8}$$

where $\mathcal{D}_{jt}^B(\cdot)$ and $\mathcal{D}_{jt}^H(\cdot)$ are the equilibrium number of Black and Hispanic households in neighborhood j. These expressions capture the fact that neighborhood demographic composition is itself an equilibrium outcome of household sorting: b_{jt} and h_{jt} enter utility through (2), but they must also equal the demographic shares implied by the resulting allocation of households across neigh-

[9] Prior evidence suggests limited impacts of neighborhood relocation on adults' earnings and employment on average (e.g., Chyn, 2018; Chyn, Collinson and Sandler, 2025). By contrast, the evidence of long-run gains from improved neighborhood exposure is substantially stronger for children (Chyn and Katz, 2021).

[10] Estimating a dynamic model with forward-looking behavior for displaced public-housing residents would require very strong assumptions given the availability of location choices in only two time periods. For the broader sample of Cook County residents, we only have access to repeated cross-sectional data, which makes estimation of a dynamic model especially difficult.

borhoods. As we show below, allowing for endogenous demographic changes has a first-order impact on the welfare consequences of demolitions.

6 Quantification of the Model

This section quantifies the model's key objects and summarizes the steps used to translate estimated primitives into equilibrium and welfare calculations. We begin by estimating household demand, separately for initial public-housing residents and for the broader Cook County population, using the identifying variation and data sources described above to recover preference parameters, moving frictions, and neighborhood mean utilities. We then calibrate housing supply, which—together with estimated demand—closes the model and allows us to solve for equilibrium prices and sorting in both the observed economy and counterfactual scenarios. With the full set of demand- and supply-side objects in hand, we summarize the resulting preference and moving-cost estimates and discuss their economic implications, and we conclude by assessing model fit and validation moments that discipline the equilibrium predictions used in our welfare and policy counterfactuals.

6.1 Demand Estimation

A necessary step to study the consequences of public housing demolitions using our model is to obtain estimates of the household preference parameters. As discussed in Section 5, households choose among neighborhoods within Cook County and an outside option $j = 0$ (living outside Cook County within Illinois). Given that the scale of utility in logit models is not identified without further assumptions, we normalize the indirect utility of the outside option to $\delta^k_{0t} = 0$.[11] This normalization implies that δ^k_{jt} only identifies utility differences with respect to the outside option (e.g., Train, 2009). This normalization also clarifies the source of identifying variation in our setting: preference parameters are disciplined by differences in choices across neighborhoods *within* Cook County.

We estimate demand-side objects separately for (i) initial public-housing residents, using administrative microdata and quasi-experimental demolition exposure, and (ii) the broader Cook County population, using tract-level Census/ACS counts and an instrumental variable approach. We present the public-housing approach first and then turn to the broader Chicago-area population.

Public Housing Residents: Our estimation treats initial public housing residents as a distinct group ($k = ph$) and leverages two key features of our analysis to identify their preferences and moving costs. First, we have administrative data that allows us to track if public housing residents move out of public housing, which we use to estimate moving costs. Second, the quasi-experimental pattern of public housing demolitions provides a source of plausibly exogenous

[11] Counterfactual simulations are identified under this normalization as long as the utility level outside Cook County does not change in response to demolitions inside Chicago (Kalouptsidi, Scott and Souza-Rodrigues, 2021).

variation in housing circumstances. The fact that displaced households are forced to relocate and face voucher-subsidized private-market rents provides credible variation to identify preferences over housing costs.

To capture the institutional details of public housing and vouchers, we specify the household-specific price term p_{ij} and two moving-cost components as functions of displacement and relocation. Because our public-housing sample tracks location choices over 1995–2000 and our richest neighborhood covariates are measured in 2000, we use year-2000 neighborhood characteristics in estimation and suppress time subscripts throughout this subsection. We allow the price of housing, p_{ij}, to vary across households as follows. If a public housing resident is not displaced by demolition, denoted by $D_i = 0$ and remains in public housing, they pay a standard cost-of-living amount c_j. If individuals are not displaced from public housing and choose to leave public housing, then they must pay the full market rent, r_j, along with the 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.[12] In addition, we include two moving costs: one for leaving the initial housing unit, MC^{ph}_{unit}, and another for leaving the initial neighborhood, MC^{ph}_{nbhd}. These parameters reflect both financial and psychological costs associated with moving, and we parametrize them so that households who leave their initial neighborhood pay both costs.

For estimation, we use a two-step approach that combines maximum likelihood estimation (MLE) in the first step and 2SLS in the second. This two-step approach allows us to address the potential endogeneity of neighborhood covariates in the second step, namely neighborhood demographics. In our first step, we use MLE to estimate $(\alpha^{ph}_p, \delta^{ph}_j, MC^{ph}_{\text{nbhd}}, MC^{ph}_{\text{unit}})$ using the individual level data. While the model implies that there is a positive probability of households moving to each neighborhood, in practice we do not observe moves to many neighborhoods, in part because the CHA data has a limited number of observations. Galiani, Murphy and Pantano (2015) confront the same challenge when estimating a location choice model based on the Moving to Opportunity (MTO) experiment. We adopt the same solution by maximizing the log-likelihood subject to the constraint that the model-implied choice probabilities match the empirical choice probabilities, $\tilde{P}_{jk}$, for Black households with income below $20,000 in the 2000 Census—a group that closely matches the demographic and income profile of initial public-housing residents and is measured with a much larger sample size in the Census.[13] A key advantage of this two-step approach is that it allows us to use instrumental variables to identify preference parameters for specific components of mean utility.

[12] In our empirical analysis, we calculate c_j as the difference between median gross rent (which includes utilities) and median contract rent (r_j, which does not) in the 2000 Census. 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.

[13] Our estimation routine recovers the mean utility, δ^{ph}_j, via a contraction mapping that, for given $(\alpha^{ph}_p, MC^{ph}_{\text{nbhd}}, MC^{ph}_{\text{unit}})$, solves for the value of δ^{ph}_j such that the model-implied choice probability matches the empirical choice probability.

In the first step, we identify the coefficients on price sensitivity and moving cost parameters by exploiting individual variation in displacement status and observed mobility behavior. A recurring challenge in the literature on neighborhood sorting is that price sensitivity is difficult to identify credibly because rents are endogenous to neighborhood demand and unobserved quality. However, in our setting, identification of the price coefficient for public housing types α_p^{PH} comes from quasi-experimental variation in exposure to public housing demolitions, which generates plausibly exogenous differences in the effective housing prices households face outside of public housing. Chyn (2018) provides supporting evidence for the credibility of this strategy by documenting no detectable differences in the baseline characteristics of those households who faced demolition-induced displacement and those who did not. Moreover, he documents that idiosyncratic maintenance problems, such as flooding due to the bursting of building pipes, appear to drive the variation in which buildings were demolished during the early demolition period. In our setting, exogenous assignment of displacement status D_i therefore acts as an as-if randomly assigned shock that changes household rental-price schedules: displaced households gain access to Section 8 vouchers and thus face lower out-of-pocket rents in private market units than otherwise similar non-displaced households. In sum, the demolition-induced shift in effective prices helps identify α_p^{ph}.

Moving-cost parameters are disciplined by distinct features of observed mobility in the administrative panel. For public-housing households, the moving cost MC_{unit}^{ph} governs the overall propensity to exit the initial public-housing unit: higher MC_{unit}^{ph} lowers the probability of leaving the initial unit holding destination utilities fixed. Conditional on leaving the unit, the neighborhood moving cost MC_{nbhd}^{ph} reduces the share of movers who leave the initial neighborhood relative to those who move within neighborhood. Because we observe both origin and destination locations, the data separately identify these components of moving frictions. This logic mirrors the dynamic neighborhood-choice literature: as in Galiani, Murphy and Pantano (2015) and Bayer et al. (2016), moving costs are disciplined by the move-stay margin and origin-destination patterns; in our setting, demolition-induced displacement and administrative panel data sharpen identification by cleanly separating frictions associated with leaving the initial unit versus leaving the initial neighborhood.

Having recovered neighborhood mean utilities from the first step, we next attribute cross-neighborhood variation in these utilities to observed neighborhood attributes using an instrumental variable (IV) strategy. Specifically, we use 2SLS to regress $\hat{\delta}_j^{ph}$ on the underlying neighborhood characteristics:

$$\hat{\delta}_j^{ph} = \alpha_b^{ph} b_j + \alpha_h^{ph} h_j + \alpha_{Inc}^{ph} \ln(Inc_j) + \alpha_{PH}^{ph} PH_j + \theta^{ph} x_j + \xi_j^{ph}, \tag{9}$$

where ξ_j^{ph} is the residual unobserved neighborhood quality component for public housing residents. The shares of households that are headed by Black or Hispanic individuals, b_j and h_j, as well as median household income, Inc_j, come from the 2000 Census. 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.[14] The vector x_j 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. All of these variables are measured for year 2000.

We treat the vector of neighborhood demographic characteristics, $(b_j, h_j, \ln(Inc_j))$, as endogenous because households could sort on the basis of unobserved neighborhood quality, ξ_j^{ph}. We instrument for those demographic characteristics using shift-share variables that predict neighborhood demographic composition based on the composition of a neighborhood's residents in 1990 and nationwide population shifts. In particular, we construct the predicted change in the population of group g in tract j between 1990 and 2000 as:

$$\Delta PredPop_j^g = \sum_b s_{b,p,1990}^g \left(Pop_{b,2000}^g - Pop_{b,1990}^g \right) f_{p,j},$$

where $s_{b,p,1990}^g$ is the share of US residents from birth state or country b that live in public-use microdata area (PUMA) p in 1990, $(Pop_{b,2000}^g - Pop_{b,1990}^g)$ is the change in the number of people from birth state/country b that reside in the entire US between 1990 and 2000, and $f_{p,j}$ is the share of land area in PUMA p that is in tract j, which maps PUMA-level variation into tract-level variation.[15] We construct $\Delta PredPop_j^g$ separately for Black individuals, Hispanic individuals, and income groups (quartiles of the nationwide income distribution). We exclude individuals born in Illinois to isolate population shifts due to migration rather than more gradual demographic aging change. These instruments leverage variation in neighborhood demographics driven by aggregate shocks to nationwide population: neighborhoods with a higher concentration of people from an origin with a faster-growing total population are predicted to receive more in-migrants than other neighborhoods. This logic follows the standard shift-share approach used to instrument for immigration-driven local demographic change across cities (e.g., Card, 2009).

Cook County Residents: For the broader Cook County population, we estimate the demand system using tract-level population counts from the Census and ACS in 1990, 2000, and 2010. 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

[14] In particular, we use data from the CHA containing the number of occupied public housing units in each tract as of January 1999. Because only around 70% of public housing units were occupied at this time (Chyn, 2018), we inflate these occupancy numbers by 1/0.7 to get an estimate of the number of public housing units.

[15] This is necessary because publicly available Census microdata only identify the PUMA where people reside.

more exposed to changes triggered by demolitions.[16] For each group-year, we construct tract-level location shares $\hat{P}^k_{jt}$ as the fraction of Illinois households of race-by-income group k residing in tract j.

Estimating tract-level location shares for small race-by-income groups using Census and ACS tabulations raises two practical challenges. First, tract-level counts—especially in the ACS—are measured with sampling error, which can introduce noise into estimated shares.[17] Second, for some tract-by-group-year cells the raw share estimates can be exactly zero, which is inconsistent with the logit structure of the model and complicates the recovery of mean utilities. To address these issues, we follow prior work (e.g., Scott, 2013) and construct smoothed location shares $\tilde{P}^k_{jt}$ by locally averaging the raw frequency shares $\hat{P}^k_{jt}$ across nearby tracts using distance-based weights:

$$\tilde{P}^k_{jt} = \sum_{n=1}^{J} w_{jn}\,\hat{P}^k_{nt},$$

where $w_{jn} \propto (1 + \text{dist}(j,n))^{-1}$ depends on the distance between tract centroids and is normalized so that $\sum_{n=1}^{J} w_{jn} = 1$ for each tract j. In what follows, we treat $\tilde{P}^k_{jt}$ as our empirical counterpart to the model-implied choice probabilities.

Because these data are repeated cross-sections rather than an individual panel, we do not observe origin-destination transitions for the broader population and therefore do not estimate the across-tract moving cost. Instead, we calibrate this cost using estimates from Bayer et al. (2016).[18] We treat this calibrated moving cost as governing the decennial move-stay margin in our setting, and we show in Section 6.4 that our main results are not sensitive to alternative calibrations.

Given the calibrated moving cost and the location shares $\tilde{P}^k_{jt}$, we recover group-year mean utilities δ^k_{jt} that rationalize observed sorting patterns across neighborhoods. Intuitively, δ^k_{jt} summarizes the average attractiveness of tract j to households in group k in year t—net of moving frictions—and is the object that we subsequently relate to rents and neighborhood characteristics.[19] In terms of estimation, we recover the vector $\{\delta^k_{jt}\}_{j=1}^{J}$ by making the model-predicted population shares at time t equal to the empirical population shares at time t for all locations j, using a contraction mapping.[20]

Having recovered $\hat{\delta}^k_{jt}$, we estimate preference parameters by relating these mean utilities to

[16] 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.

[17] 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 (roughly 1 percent in each year).

[18] We use the results from Table 4 in Bayer et al. (2016) along with the average value of housing in Chicago. We convert their annual magnitude into a ten-year present value calculation to reflect the ten-year period between our census samples. The resulting moving cost is $23,000.

[19] Relative to equation (2), we abuse notation slightly for the Cook County resident model by including $\alpha^k_p \ln(p_{jt})$ in the mean utility, δ^k_{jt}.

[20] Specifically, for each group k and period $t = 2000, 2010$, we choose $\{\delta^k_{jt}\}_{j=1}^{J}$ so that the following relationship holds:

$$\tilde{P}^k_{jt} = \sum_{o} \tilde{P}^k_{ot-1} \frac{\exp(\delta^k_{jt} - MC \cdot \mathbb{1}\{j \neq o\})}{\sum_{j'} \exp(\delta^k_{jt} - MC \cdot \mathbb{1}\{j' \neq o\})},$$

observed neighborhood characteristics using a fixed-effects model that relies on aggregate repeated cross sections. Specifically, we consider the following specification:

$$\hat{\delta}^k_{jt} = \alpha^k_p \ln(p_{jt}) + \alpha^k_b b_{jt} + \alpha^k_h h_{jt} + \alpha^k_{Inc} \ln(Inc_{jt}) + \alpha^k_{PH} PH_{jt} + \theta^k x_{jt} + \lambda^k_j + \lambda^k_t + \tilde{\xi}^k_{jt}, \tag{10}$$

where λ^k_j are tract fixed effects that represent time-invariant neighborhood attributes (e.g., distance to the central business district) and λ^k_t are year fixed effects that capture shocks to all Cook County tracts relative to the outside option.[21] Both sets of fixed effects are allowed to vary flexibly by race-by-income group k. The right-hand-side variables mirror those used for public housing residents. We measure p_{jt} using median gross rents (contract rent plus utilities) from the Census/ACS; unlike in our public housing approach, these households do not receive vouchers, so out-of-pocket costs do not vary at the household level within a tract-year. We measure neighborhood demographics and median income using the same sources. We measure the public-housing stock using CHA data on the baseline stock and subsequent demolitions.[22]

Given that equation (10) includes tract and year fixed effects, the remaining identification concern for our approach is tract-specific time-varying unobserved quality $\tilde{\xi}^k_{jt}$ that may be correlated with within-tract changes in rents and demographics. We address this concern by estimating (10) using a 2SLS approach. In particular, we treat the vector $(\ln(p_{jt}), b_{jt}, h_{jt}, \ln(Inc_{jt}))$ as endogenous and instrument for these variables using two families of instruments. First, we use shift-share predictors of demographic change that combine baseline settlement patterns with nationwide population shifts by origin group, analogous to the instruments used for public-housing residents. Second, to address the endogeneity of rents, we use instruments based on the 1923 Chicago zoning ordinance (Shertzer, Twinam and Walsh, 2016) interacted with year indicators, which generate differential predicted rent growth across tracts over time.[23]

We implement this IV strategy using a three-step procedure that follows the optimal-instrument literature (e.g., Chamberlain, 1987; Reynaert and Verboven, 2014) and has been adopted in neighborhood-sorting applications (Bayer, Ferreira and McMillan, 2007; Wong, 2013; Calder-Wang, 2019; Anagol, Ferreira and Rexer, 2021; Allen, Arkolakis and Takahashi, 2020). In the first step, we obtain preliminary 2SLS estimates from equation (10) using our baseline set of instruments. In the second step, we solve for the model-implied equilibrium rents and location choices given these preliminary estimates under the normalization that there are no unobserved time-varying demand shocks (i.e.,

where the left-hand-side of the previous equation is the observed probabilities for tract j at time t, and the right-hand-side are model-predicted probabilities of destination j summed over the origin distribution at time $t-1$. We solve the resulting system of equations using a contraction mapping following (Berry, Levinsohn and Pakes, 1995).

[21]Note that shocks common across all tracts in Illinois, such as changes in inflation rates, are differenced out through the normalization that $\delta^k_{0t} = 0$.

[22]Data from the CHA provide us with information about the number of public housing units demolished in each period. We do not have data directly measuring public housing occupancy or the stock of public housing units beyond 2000. 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. The time-invariant 1990 level is absorbed by tract fixed effects, so identification comes from differential changes in public-housing exposure over time.

[23]In particular, we construct the share of land in each census tract zoned for commercial use, manufacturing use, and volume districts 1–3 (which restricted building height and land area coverage). We interact these variables with year fixed effects to leverage variation over time that is predicted by these historical factors.

$\tilde{\xi}^k_{jt} = 0$), and we use the resulting simulated allocation to construct predicted rents and predicted demographic shares for each tract and year. In the third step, we re-estimate equation (10) using these simulated rents and simulated demographic shares as instruments for the endogenous variables. Intuitively, this procedure concentrates variation from many underlying instruments into functions that are especially relevant for the endogenous regressors of interest and incorporates information from the equilibrium conditions in (6)–(8).

The validity of this IV approach relies on the assumption that our instruments are not systematically related to changes over time in unobserved neighborhood quality, conditional on the included controls. Figure 5 presents a placebo test of the associated pre-trend condition by relating 1990–2000 changes in rents and house values to (i) subsequent demolition intensity and (ii) our instrument-predicted rent changes from 2000–2010. Consistent with our exclusion restriction, pre-period rent growth is unrelated to both measures (Panels A and B), and the same holds for house values (Panels C and D).

6.2 *Supply Calibration*

In addition to estimating household demand, our counterfactual exercises require a specification for housing supply. As discussed in Section 5, we assume an isoelastic supply relationship, $\mathcal{S}_{jt}(p_{jt}) = \kappa_{jt} p_{jt}^{\psi}$, where ψ governs the elasticity of housing supply and κ_{jt} is a tract-by-year supply shifter.

We calibrate the supply elasticity ψ using tract-level estimates for Chicago from Baum-Snow and Han (2024), who report unit supply elasticities between 0.106 and 0.220. In our baseline analysis we set $\psi = 0.183$, the mean value.[24] Given ψ, we recover the supply shifter κ_{jt} by combining (5) with baseline equilibrium quantities implied by estimated demand. Specifically, we compute the model-implied demand for each tract-year under the estimated preference parameters and setting unobserved demand shocks to zero, $\tilde{\xi}^k_{jt} = 0$, and then choose κ_{jt} to rationalize observed prices:

$$\hat{\kappa}_{jt} = \mathcal{D}_{jt}(\mathbf{p}_t, \mathbf{b}_t, \mathbf{h}_t, \mathbf{x}_t, \mathbf{0}; \hat{\alpha}, \widehat{MC}) \big/ p_{jt}^{\psi}. \tag{11}$$

This approach ensures that the calibrated supply shifters are not mechanically contaminated by time-varying unobserved demand components.[25]

6.3 *Equilibrium and Welfare Accounting*

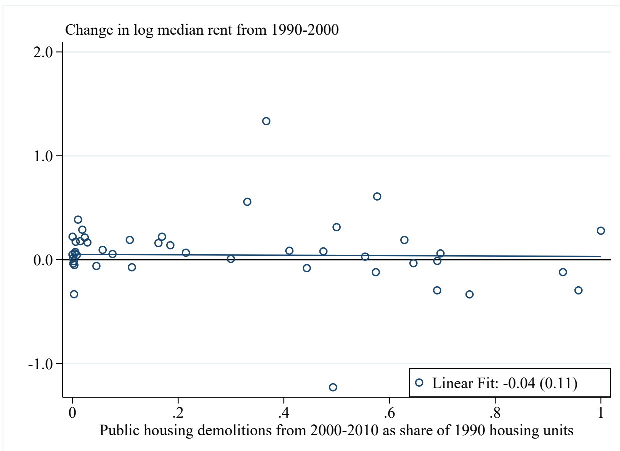
With the estimated demand system and calibrated housing supply in hand, we use the model to compute equilibrium outcomes under observed and counterfactual policy environments. These equilibrium computations serve two purposes in the paper. Most importantly, they deliver the

[24]The in-sample fit of the model reported below is considerably worse when using tract-specific housing supply elasticity estimates from Baum-Snow and Han (2024), possibly because each of these 1,240 parameters is estimated with some error that is not incorporated into our model.

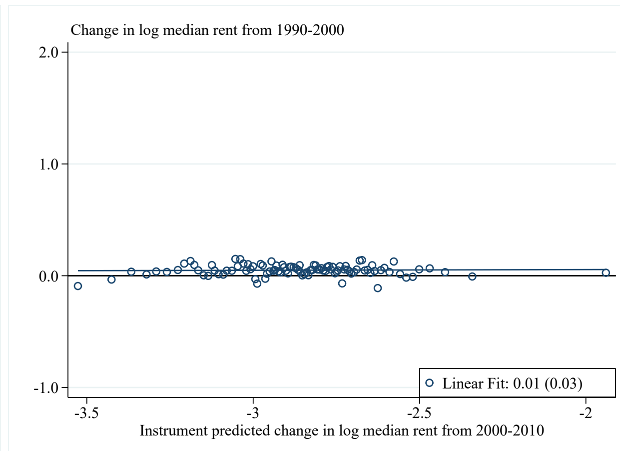
[25]The correlation between observed tract-level quantities from the Census/ACS and the model-implied quantities used in the calibration is 0.98. Our welfare results remain virtually unchanged if we instead use observed quantities to calibrate θ_{jt}.

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

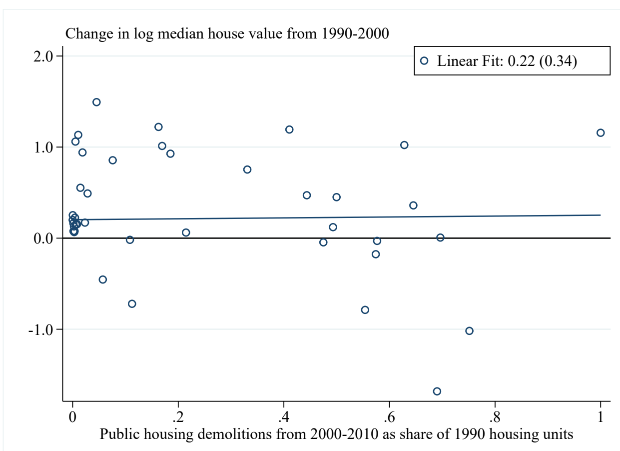
(a) Change in Log Rent Against Demolitions



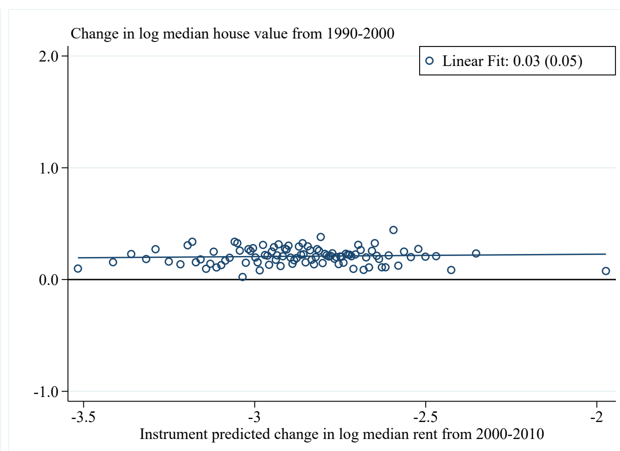
(b) Change in Log Rent Against Instrument



(c) Change in Log House Value Against Demolitions



(d) Change in Log House Value Against Instrument



Notes: This figure plots the change in log median rent and log median house value from 1990–2000 against the number of public housing units demolished from 2000–2010 as a share of the number of occupied housing units in 1990 (Panels A and C) and the predicted change in log rent from 2000–2010 based on our instrumental variable procedure (Panels B and D). In Panels A and C, each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition. In Panels B and D, each dot represents the average change for each percentile of the instrument-predicted change in log rent.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

equilibrium objects used in the welfare and policy counterfactual results in Section 7. They also underlie the validation exercises in Section 6.5, which assess how much cross-tract and over-time variation in rents is explained by the explicitly modeled demand, sorting, and supply components.

For any tract-level environment—defined by the public-housing stock, any policy-induced price schedules faced by displaced households, and the set of observed neighborhood attributes—we solve for the equilibrium vector of rents and neighborhood composition that clears housing markets and is consistent with demographic sorting. We compute an equilibrium using the mapping described in Section 5.2 and denote the solution as $(\mathbf{p}(\mathcal{E}), \mathbf{b}(\mathcal{E}), \mathbf{h}(\mathcal{E}))$ for environment $\mathcal{E}$. Unless otherwise noted, we set the time-varying unobserved component of neighborhood quality to its unconditional mean (i.e., $\tilde{\xi}^k_{jt} = 0$ for all j, t, and k) when computing equilibrium outcomes. This restriction isolates the explanatory power of observed neighborhood attributes and the model's structural components; Appendix B provides details on the numerical procedure.[26]

Our baseline comparison is the observed post-demolition environment versus a no-demolition counterfactual. In the observed environment, tracts reflect the realized reduction in the public-housing stock. In the no-demolition counterfactual, we restore tract-level public-housing stock to its pre-demolition level (held at the year 2000 shares) and remove demolition-induced displacement, holding other neighborhood characteristics fixed at their observed values in a given outcome year. This counterfactual is meant to mirror the institutional baseline in Section 2: the Plan for Transformation was authorized around 2000 and subsequent years saw the major demolition-and-relocation wave, so the year-2000 stock provides a natural benchmark for "pre-Plan" public-housing capacity in our equilibrium comparisons. In our simulations, we also account for the fact that demolitions shift public-housing residents into the private housing market, which could create extra pressure on rents and generate "pecuniary externalities" of public housing demolitions.

Given equilibrium outcomes for each scenario, we measure welfare in rent-equivalent units for residents of Cook County. Using the estimated preference parameters for each group group k, 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 group-specific average renter consumer surplus in closed-form solution as follows:

$$\mathcal{CS}^k_t(\mathbf{p}^k, \mathbf{b}, \mathbf{h}, \mathbf{x}; \alpha^k, MC^k) = \sum_{\ell=0}^{J} s^k_{\ell t-1} \log\left(\sum_{j \in \mathcal{J}^k} \exp\left(v^k_{jt|\ell}(\mathbf{p}^k, \mathbf{b}, \mathbf{h}, \mathbf{x}; \alpha^k, MC^k)\right)\right) + \Gamma,$$

where the indirect utility associated with moving from ℓ to j for a household of type k is $v^k_{jt|\ell} \equiv \delta^k_{jt}(\cdot) - \alpha^k \ln(p^k_{jt}) - MC^k_{\ell,j}$, origin locations are indexed by $\ell \in \{0, \ldots, J\}$, the set of potential destinations—excluding the outside option—for type k is given by $\mathcal{J}^k$, and the share of type k households living in location ℓ in period $t-1$ is $s^k_{\ell t-1}$. The destination set, $\mathcal{J}^k$, is the same for all of the broader Cook County population groups. In some counterfactuals, public-housing residents

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

have one additional choice: staying in their original public-housing unit. Note that effective rents are type-dependent because displaced households may face voucher price schedules. We therefore write p_{jt}^k for the out-of-pocket rent faced by group k in tract j at time t; for non-voucher households p_{jt}^k equals the market rent, while for voucher recipients it incorporates the voucher payment rule. Within each group k, we apply the same rent-based welfare measure to all households. For renters, welfare changes are captured entirely by changes in rent-equivalent consumer surplus. For homeowners, we use the same rent-equivalent measure for the consumption value of location choices and additionally account for the change in rental income implied by equilibrium rent changes (under our portfolio assumption below).

To express welfare changes in monetary terms, we define the group-specific rent equivalent for moving from scenario 0 to scenario 1 as a change in group-k consumer surplus divided by the marginal utility of housing consumption, following Small and Rosen (1981):

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

where μ^k is the marginal utility of housing consumption for group k.[27] Note that positive values of the rent equivalent indicate higher welfare in the counterfactual world in scenario 1 versus 0.

The rent-equivalent welfare measure above is naturally interpreted for renters, for whom changes in equilibrium rents represent out-of-pocket payments. For homeowners, however, rents reduce utility, but higher equilibrium rents also increase the income earned on housing assets. In our accounting, we therefore add the implied change in rental income to the rent-equivalent welfare change for homeowners—i.e., we treat rent payments as internal transfers for owners, rather than pure resource losses.

For simplicity, we assume homeowners hold a fully diversified portfolio of housing, so that all homeowners receive the same change in rental income, equal to the average rent change across tracts in Cook County, $\Delta\bar{r}$. Under this assumption, the welfare change for group k is

$$\Delta W^k \; = \; RE^k \; + \; s_{\text{home}}^k \cdot \Delta\bar{r},$$

where s_{home}^k is the share of group-k households who are homeowners and RE^k denotes the rent-equivalent change implied by the location-choice model.

6.4 *Housing Demand Results*

Table 2 reports estimates of the preference parameters. Column 1 reports results for the initial public housing residents, while columns 2–7 report results for the broader population of

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

$$\mu^k \equiv -\frac{\alpha_p^k}{\sum_j \frac{N_j^k}{N^k} p_j},$$

where the numerator is the estimated preference parameter for housing costs and the denominator is the choice-probability-weighted average of neighborhood rents.

Cook County.[28] As expected, we find that all households have a distaste for 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 consistently a disamenity. Notably, a comparison of the IV-based results and OLS results for Cook County residents in Appendix Table A.3 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.

Table 2: Estimates of Neighborhood Preference Parameters

		Cook County Residents					
	Public Housing Residents (1)	Poor Non-Hispanic White (2)	Poor Black (3)	Poor Hispanic (4)	Non-Poor Non-Hispanic White (5)	Non-Poor Black (6)	Non-Poor Hispanic (7)
Log median rent	-0.649***	-0.446***	-0.0602**	-0.272***	-0.0646***	-0.0655***	-0.271***
	(0.039)	(0.0717)	(0.0244)	(0.0520)	(0.0133)	(0.0158)	(0.0492)
Black share	0.898***	-0.134	0.342***	0.287***	-0.114***	0.174***	0.0975*
	(0.072)	(0.0817)	(0.0223)	(0.0563)	(0.0147)	(0.0222)	(0.0569)
Hispanic share	0.467***	-0.0623	0.0916***	0.306***	-0.137***	0.0638***	0.151***
	(0.092)	(0.0454)	(0.0177)	(0.0367)	(0.00886)	(0.0104)	(0.0332)
Log median HH income	0.277***	0.0841***	0.0203***	0.0291*	0.0176***	0.0107**	0.0237
	(0.098)	(0.0225)	(0.00654)	(0.0155)	(0.00380)	(0.00445)	(0.0152)
PH as share of all housing	-0.710***	-0.439***	-0.0591*	-0.289***	-0.0811***	-0.114***	-0.316***
	(0.092)	(0.111)	(0.0352)	(0.0801)	(0.0184)	(0.0257)	(0.0781)

Notes: Table presents estimates of preference parameters for public-housing residents (column 1) and the broader set of Cook County residents (columns 2–7). The results in column 1 are based on equation (9) and an approach that uses MLE and 2SLS on administrative data tracking the location of public-housing residents. The estimates in columns 2–7 are based on equation (10) and an approach that uses 2SLS on census/ACS data. We estimate preferences separately for each group. Standard errors in columns 2–7 are clustered at the tract level to account for the use of repeated cross-section data.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

For public housing residents, we estimate substantial utility costs from moving. In rent-equivalent terms, the estimates imply that the present value cost (over a ten-year interval) of moving out of public housing is equivalent to \$16,784, and the corresponding cost of moving to a different tract is \$36,287 (i.e., a total cost of moving is \$53,071).[29] These magnitudes are broadly

[28]We present the main results for non-Hispanic White, Black and Hispanic households for clarity. Appendix Tables A.5 and A.7 report 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.

[29]We treat the moves observed in the administrative data as occurring over a one-year period for the purpose of estimating the model. For the estimates based on census/ACS data, we consider moves over a ten-year horizon (from 2000 to 2010). To express one-year moving costs on this decadal timeline, we assume the relevant relocation occurs at the end of the ten-year interval and discount back using an annual discount factor of 0.95: $C^{10} = 0.95^{10}C^1 \approx 0.60\,C^1$ for cost C. This timing convention is conservative; if moves occur earlier within the decade, the implied present value

in line with dynamic models of housing and neighborhood choice. For example, the estimates in Bayer et al. (2016) imply a moving cost of $23,000 in ten-year present value. Our estimated moving cost for public-housing residents is larger, consistent with large mobility frictions for low-income households living in subsidzed, long-tenure public housing units. In a distinct but related setting, Almagro and Domínguez-Iino (2024) find that moving costs are almost four times larger for low-income individuals than high-income individuals.

These preference estimates can be used to calculate the implied willingness to pay for different neighborhood characteristics in terms of annual rents.[30] For example, initial public housing residents are willing to pay $148 in annual rents to reduce the share of the housing stock that is public housing by 1 percentage point and $185 higher rents to increase the Black population share by 1 percentage point.

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 similar to the willingness to pay for same-race neighbors for initial public housing residents.

Our estimation of preferences for the broader Cook County population relies on instrumental variables to a greater degree than our estimation for public-housing residents, which raises natural questions about robustness. Appendix Tables A.4–A.7 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, 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 1990–2000 changes in neighborhood characteristics. Fourth, we estimate a model on a sample that excludes all census tracts within one mile of the Cabrini-Green public housing project (column 5) due to concern that unobserved trends in gentrification may confound our ability to obtain unbiased estimates. Fifth, we use replace the income-based

would be larger.

[30] Coefficients in Table 2 are measured in utils and thus cannot be compared across groups.

shift-share instruments with education-based variables (column 6). Finally, we add as instruments the average depth to bedrock and average slope of land—factors that can influence construction costs (column 7). The results from Appendix Tables A.4–A.7 show that the estimated preference parameters are quite robust.

Our estimated preferences for the broader Cook County population also rely on the calibrated moving cost, which is used to convert choice probabilities into mean indirect utilities. In Appendix Tables A.8 and A.9, we show that preference estimates are similar when scaling the moving cost down to a range of different values, including when moving costs are assumed to be zero. While moving costs are important for understanding the welfare consequences of public housing demolitions, as shown below, they do not drive the key ways in which households trade off neighborhood characteristics.

6.5 *Assessing Model Fit*

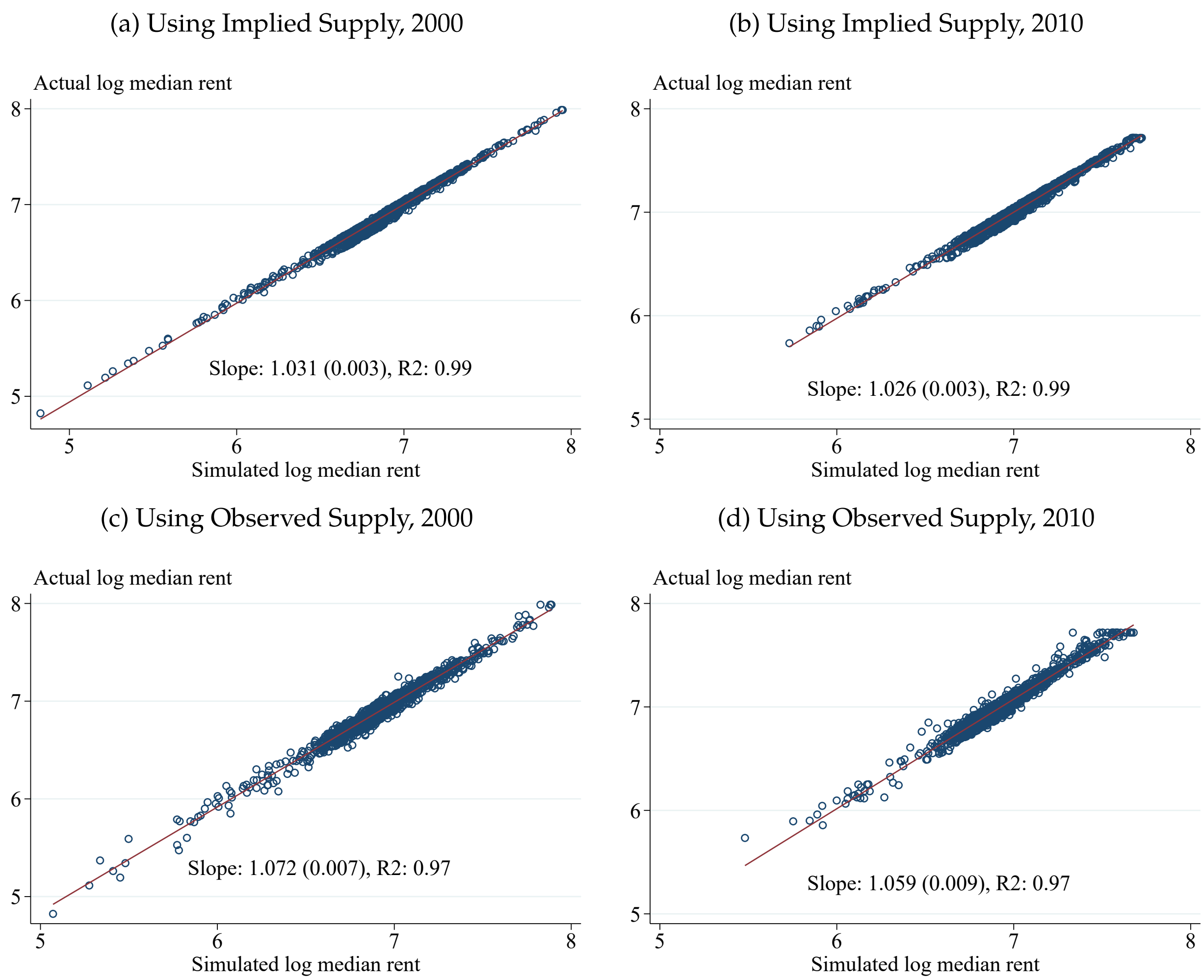
Before describing the welfare consequences of public housing demolitions, we conduct one in-sample and one out-of-sample validation exercise to assess how well our model fits equilibrium rental prices. In both exercises, we focus on the explanatory power of the explicitly-modeled elements by setting the time-varying unobserved component of neighborhood quality equal to zero (i.e., $\tilde{\xi}_{jt}^{k} = 0$ for all j, t, and k). This restriction asks how much cross-tract and over-time variation in rents can be accounted for by the demand system, endogenous sorting, and housing supply, using only observed neighborhood attributes.[31] 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 6.2. 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 2010. These results demonstrate a high degree of in-sample fit due to observable attributes in the data.

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

[31] Allowing $\tilde{\xi}_{jt}^{k}$ to absorb the remaining demand residual would mechanically deliver a perfect in-sample fit.

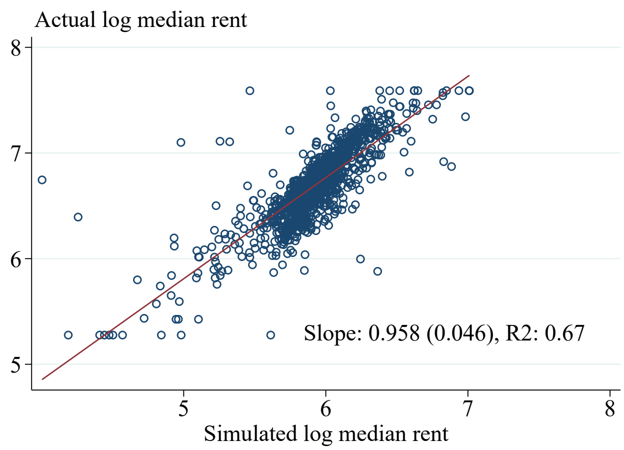
Figure 6: Assessing In-Sample Fit of Structural Model Using Rent Data



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.

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 (6)–(8) allows us to solve for the endogenous variables in this exercise. Importantly, we do not use any data from 1990 on the endogenous variables in this procedure. Our test is a comparison of the resulting equilibrium rents simulated out-of-sample for 1990 against the actual rents. The results in Figure 7 show that there is an almost one-to-one relationship between actual and simulated rents on average. Moreover, the simulated rents explain 67 percent of the cross-tract variation in actual rents. This out-of-sample validation exercise underscores the strong fit of the model.

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 (6)–(8).
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

7 Impacts of Public Housing Demolitions

What impact did public housing demolitions have on welfare? To answer this, we use our estimated demand parameters and our calibrated housing supply with the counterfactual and welfare accounting procedures described in Section 6. Throughout, we distinguish the direct consequences for public housing residents—forced relocation and voucher assistance—from the general-equilibrium effects on other households operating through endogenous rents and neighborhood composition, with equilibrium rent increases playing a central role in the distribution of gains and losses. We begin by presenting the headline welfare results and their incidence across groups, and then we use decomposition exercises to clarify the mechanisms driving these effects. We also describe how the implied equilibrium impacts vary across neighborhoods throughout Chicago. Finally, we assess robustness to key modeling assumptions in Section 7.4, including the possibility of multiple equilibria and allowing for intensive margin adjustments in housing consumption.

7.1 Welfare Impacts

Table 3 reports the welfare changes in rent-equivalent units for the average household due to public housing demolitions for public housing households and the broader set of Cook County race-by-income groups. These estimates compare the observed 2010 equilibrium to a 2010 no-demolition counterfactual constructed as described in Section 6.3, in which the tract-level public-housing stock is restored to its pre-demolition level (held at 2000 shares) and demolition-induced displacement and voucher price schedules are removed, while all other exogenous neighborhood characteristics are held fixed at their observed 2010 values. For each scenario, we solve for the equilibrium vector of rents and neighborhood composition, so the reported welfare changes incorporate both the direct effects on displaced public-housing residents and the broader general-equilibrium response for other households operating through endogenous rent and sorting adjustments. Because the comparison is 2010-to-2010, the results are not mechanically driven by aggregate changes from 2000 to 2010 in Cook County relative to the outside option.

The results show that there are large, negative effects for public housing residents, who see a decrease in utility amounting to \$6,582 in rental equivalent units. The magnitude is also consistent with our preference estimates, which imply substantial moving frictions; in this setting, demolitions compel moves that households would not otherwise choose, so the net welfare change reflects a tradeoff between the large disutility of forced relocation and offsetting gains from voucher-enabled choice (and any accompanying neighborhood and equilibrium adjustments). We unpack these mechanisms formally in the decomposition analysis below by separating the forced-relocation component from the role of voucher assistance and related equilibrium channels.

For the broader population, the pattern is sharply redistributive. Poor households—especially Black and Hispanic residents—experience welfare losses on average, while non-poor households of all races and ethnicities are made better off. A natural interpretation is that demolitions raise equilibrium rents, which reduces welfare for renters—who disproportionately consist of lower-income households. As we show below, rent increases are a central part of the general-equilibrium response. More broadly, the incidence pattern is consistent with a broader place-based policy literature emphasizing that locally-targeted improvements can be capitalized into housing prices and therefore shift gains toward groups with higher incomes and higher homeownership rates (e.g., Busso, Gregory and Kline, 2013). In population-weighted terms, the average effect is a small loss of approximately \$3 per household because non-poor households account for 80% of the total population. We also report the aggregate change in welfare for each group, which is the average change multiplied by the number of households. In aggregate, the welfare change amounts to a total welfare loss of \$4.6 million. This is driven by the large impact on public housing residents, which totals to \$92 million, while non-poor White households have a gain of \$98.1 million.[32]

[32]Given that the largest gains accrue to groups that are both numerous and have high homeownership rates, this pattern is consistent with a political-economy environment in which large-scale demolition and redevelopment can be supported even when aggregate welfare is close to zero or slightly negative.

Summing across income groups, White households gained \$95.6 million and Hispanic households gained \$4.7 million, while Black households lost \$12.8 million.

Table 3: The Overall Welfare Effects of Public Housing Demolitions

	Total households (1)	Average welfare change (2)	Aggregate welfare change, millions (3)
Public housing residents	13,985	-6,581.7	-92.0
Cook County residents			
Non-Hispanic White, Poor	120,840	-20.7	-2.5
Non-Hispanic White, Non-poor	786,279	124.7	98.1
Black, Poor	142,858	-215.9	-30.8
Black, Non-poor	300,190	60.1	18.0
Hispanic, Poor	57,177	-85.8	-4.9
Hispanic, Non-poor	254,458	37.6	9.6
Total	1,675,787	-2.8	-4.6

Notes: This table reports the aggregate rent equivalent change in welfare from the actual state of the world in 2010 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. Column 2 reports the average welfare change for households of the indicated type, and column 3 reports the aggregate welfare change in millions of dollars, which is found by multiplying the average change in column 2 by the total number of households in column 1 and then converting to millions. 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.

7.2 Welfare Decomposition and Channels

To understand the impact of demolitions on public housing residents' welfare, Table 4 provides decomposition results. This exercise selectively highlights the quantitative importance of various channels that are embedded in our model. We start in the counterfactual 2010 scenario without demolitions and turn consider a sequence of counterfactuals that "turn on" additional margins of of adjustment. Each row reports the level of average rent-equivalent welfare change relative to the no-demolition equilibrium under the scenario described by the row label. For interpretation, the marginal contribution of a given margin of adjustment is the difference between adjacent rows, since moving from one row to the next adds exactly one new channel.[33] The early rows therefore correspond to partial-equilibrium exercises in which only a subset of outcomes are allowed to adjust, while later rows incorporate displacement and the post-demolition rent and

[33]For example, the change in rent equivalent welfare shown in the first row of Table 4 is computed as:

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

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

voucher regime faced by incumbent public-housing households.

We begin by isolating "neighborhood-side" channels—changes in the local environment induced by demolitions—abstracting from the forced relocation of incumbent public housing households (though households may still change their location if it is optimal to do so). The first row reports results where the counterfactual considered is one in which public housing is destroyed but racial demographics and rents are fixed at their counterfactual 2010 levels with public housing. Because public housing residents view public housing buildings as a disamenity, the rent equivalent number from destroying public housing in the first row is 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.[34] This mechanical change in neighborhood composition decreases the utility of public housing residents, 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. These results show that demolitions disrupted areas that had a favorable demographic composition for public housing residents, but equilibrium re-sorting allows minority groups to partially recreate the demographic landscape of the disrupted communities. Finally, the fourth row allows market rents to adjust in equilibrium. As we detail further below, demolitions increase rents—for the city as a whole, average rents rise by about 2.6 percent—and this reduces rent-equivalent welfare for renters. Taken together, these channels raise the welfare of public housing residents by an amount equal to a $321 decrease in annual rents.

We next incorporate the "displacement-side" channels: the forced relocation and the post-demolition rent and voucher regime faced by incumbent public-housing households. The fifth row imposes displacement and the associated utility cost of moving out of their housing unit; this generates an average welfare loss of $16,463. 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.[35] Next, the scenario in the sixth row forces public housing households to pay market rents, which lowers their welfare by $1,742. Finally, the last row shows the incremental impact of receiving a housing voucher that can be used anywhere in the city; this raises their welfare by $11,624. 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 $6,582. The key forces that affect the welfare of displaced public housing residents are the utility

[34] 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). See Appendix Table A.2.

[35] By contrast, households only pay the neighborhood-level moving cost if they decide that moving to another neighborhood is optimal. Given the size of the moving cost, this tends to occur when households get a particularly positive idiosyncratic preference draw in another neighborhood.

Table 4: The Welfare Effects of Public Housing Demolitions for Public Housing Residents

Change in average welfare from no-demolition state	
Destroy buildings in tract	796.1
... and change neighborhood composition via demolitions	-120.0
... and change neighborhood composition via resorting	460.1
... and change housing prices (all channels faced by non-public-housing renters)	321.1
... and forced to move out of housing unit	-16,463.3
... and forced to pay market rents	-18,205.6
... and provided with housing voucher (all channels)	-6,581.7

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. 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.

Table 5: The Welfare Effects of Public Housing Demolitions for Census/ACS Sample

	Change in average welfare from no-demolition state					
	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	147	183	173	274	171	174
... and change neighborhood composition via demolitions	164	259	-190	117	120	160
... and change neighborhood composition via resorting	152	198	73	235	158	171
... and change housing prices (all channels for renters)	-182	-131	-276	-110	-178	-161
Panel B. Results for homeowners						
... and redistribute rents to homeowners (all channels for owners)	132	184	39	205	136	154
Panel C. Full equilibrium results						
Average welfare change across renters & owners	-21	125	-216	60	-86	38
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.

costs of being displaced and the benefit of gaining access to a Section 8 voucher.

For the broader Census/ACS sample of Cook County residents, Table 5 conducts a similar welfare decomposition exercise. Panel A focuses on renters and reports the same sequence of neighborhood-side counterfactuals: removing public housing as a neighborhood characteristic, allowing demographic shares to adjust mechanically with removals, allowing full re-sorting, and then allowing rents to adjust. The qualitative patterns are the same for the destruction of public housing, which is a disamenity for all groups. However, the change in demographic composition caused by demolitions has different effects on White residents, who prefer to live in neighborhoods with fewer Black residents, versus Black and Hispanic households. Once rent adjustments are included, renter households from all groups are worse off, reflecting the fact demolitions increase equilibrium housing market prices.

Finally, the next results in Table 5 summarize welfare when all endogenous channels operate for homeowners (Panel B) and the population as a whole (Panel C). The results for homeowners alone show that all racial groups have gains on average, reflecting the associated increases in equilibrium rental income. The full population results in Panel C demonstrate that there are notable disparities in the effects of demolitions. Since higher rents reduce renter welfare but increase it for homeowners, groups with higher ownership rates will have gains. Consistent with national trends, homeownership rates vary substantially across the groups in our setting. For example, in 2010, the year of our counterfactual, homeownership rates vary from 81 percent for non-poor White households to 19 percent for poor Black households.[36] As a result, welfare effects of demolitions differ sharply by race even within income groups: the average effect is -\$21 for poor White households, compared with -\$216 for poor Black households.

7.3 *Equilibrium Housing Prices and Sorting*

This section focuses on the two endogenous objects that drive welfare incidence: housing market rents and demographic sorting. We first summarize how demolitions shift equilibrium rents on average and decompose the aggregate price effect into direct increases in demolition tracts and spillovers across the broader city. We then report the estimated impacts of demolitions on net population flows and changes in demographic composition in Cook County. Finally, we document how these aggregate shifts manifest spatially, showing that neighborhoods closest to demolition sites experienced the largest rent increases and exhibited systematic re-sorting, with higher-income and White households disproportionately replacing poorer Black households in affected areas.

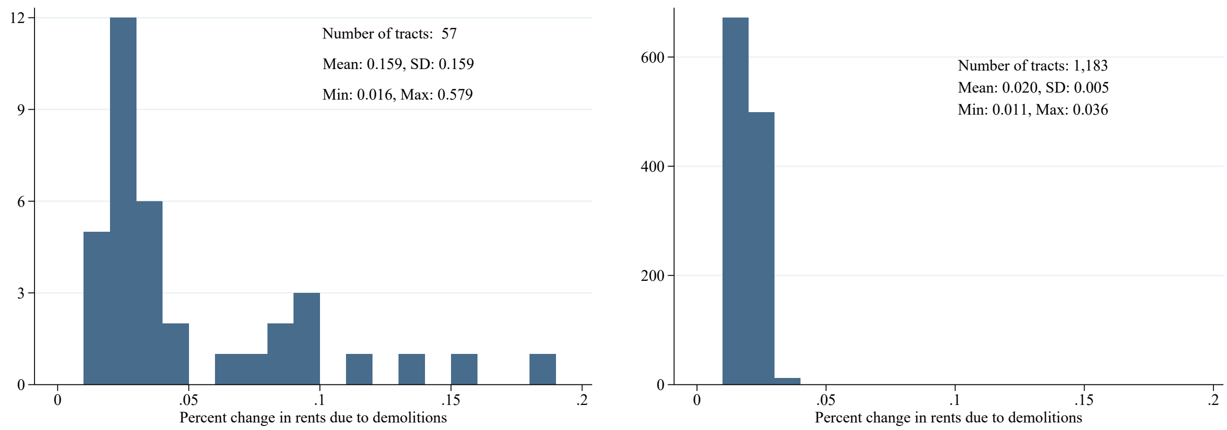
As previewed in our decomposition results above, we find that there are important equilibrium shifts in housing market rents due to public housing demolitions. Comparing the equilibrium price levels under the demolition and no-demolition scenarios, we find that average rents rise by about 2.6 percent citywide (see Appendix Table A.10). This rent response includes any direct impacts in areas targeted for demolitions and broader equilibrium spillovers as households re-optimize their

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

locations throughout the city. These spillovers are central for interpreting the welfare results: even if demolitions are geographically concentrated, a citywide upward shift in the rent schedule affects a much larger set of renters and owners and shape the incidence of demolitions across groups.

Figure 8: Distribution of Tract-Level Rent Changes Due to Public Housing Demolitions

(a) Tracts with Public Housing Demolitions (b) Tracts without Public Housing Demolitions

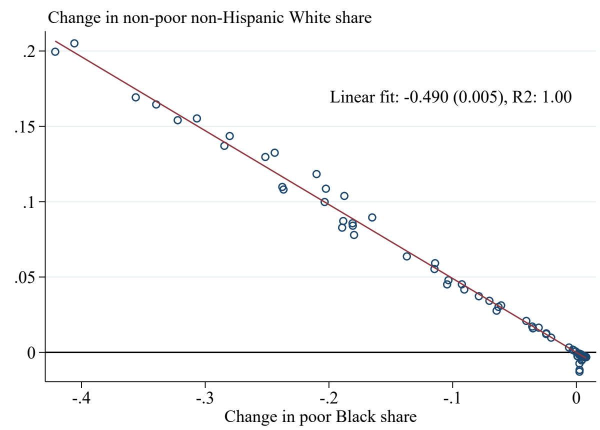


Notes: This figure displays the distribution of the percent change in median rents due to public housing demolitions. We construct this change using estimates from the model and a comparison of differences between the actual scenario in 2010 (after demolitions occurred) and a counterfactual scenario 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 18 tracts where the change exceeds the included range. 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. *Source*: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

To estimate the direct and spillover effects of public housing demolitions, we compare tract-level variables in the demolition equilibrium to those in the no-demolition counterfactual equilibrium. Figure 8 provides separate histograms to quantify how the citywide rent increase is distributed across tracts with demolitions versus the broader tracts without demolitions. Tracts with a demolition saw an average rent increase of 15.9 percent (Panel A). The distribution of changes for these tracts exhibits a fat right tail, with the largest changes occurring in neighborhoods that experienced the most extensive demolition activity. The 15.9 percent effect for tracts with demolitions is almost eight times larger than the average rent increase of 2.0 percent in tracts without demolitions (Panel B). Despite being smaller, the spillover increase is central because it accounts for 71 percent of the citywide rent response (Appendix Table A.10). The fact that prices increase even in neighborhoods without demolitions demonstrates that the effects of demolitions extend beyond treated areas: as rents rise in demolition tracts, demand shifts toward other neighborhoods, putting upward pressure on rents throughout the broader market. 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.

Having characterized the equilibrium price response, we next turn to the second endogenous force shaping incidence: demographic sorting. In the model, demolitions affect neighborhood composition through two related margins: net population flows into and out of Cook County and re-sorting across neighborhoods among households that remain in the market. Consistent with the citywide rent increase documented above, demolitions make Cook County more attractive at the margin and lead to a net increase in the number of resident households: the total number of households in the housing market rise by 8,000 (0.4 percent) relative to the no-demolition counterfactual. Demolitions also shift the demographic composition of the county: the share of households that are poor and Black declines by 4.1 percent.

Figure 9: 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.

Figure 9 summarizes the spatial incidence of demolition-induced re-sorting by plotting 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 in the poor Black household share experienced a 0.5 percent increase in the non-poor White household share. Consistent with the rent results above, these compositional changes are tightly linked to local price adjustments: across tracts, the demolition-induced percent change in rents is strongly negatively correlated with the change in the poor Black share ($\rho = -0.71$) and positively correlated with the change in non-poor White share ($\rho = 0.68$). This replacement pattern is consistent with the descriptive evidence in Figure 4, which shows that demolition exposure is systematically associated with shifts in neighborhood demographic composition.

7.4 Robustness to Model and Calibration Choices

This section assesses the sensitivity of our counterfactual welfare and incidence results to a set of key modeling and calibration choices. We consider three exercises. First, because the across-tract moving cost for the broader Cook County population is calibrated (rather than estimated from transitions), we vary this parameter over a wide range. Second, because endogenous demographic sorting can in principle generate multiple equilibria, we implement an extensive equilibrium-search procedure using many starting values to confirm that the algorithm converges to the same fixed point. Finally, we allow for an intensive margin in housing consumption by introducing Cobb-Douglas preferences (rather than assuming fixed housing consumption). We discuss each check in detail below.

Varying the Calibrated Moving Cost: Our welfare results for the broader Cook County population rely on the moving cost that we calibrate from Bayer et al. (2016). Because this parameter is imported from prior work rather than estimated directly in our setting, it is a natural object for sensitivity analysis. Accordingly, we conduct a sensitivity test that varies the calibrated moving cost governing relocation frictions for the non-public-housing population and re-estimates preference parameters and equilibrium objects, including welfare.

In theory, a reduction of moving costs makes households more responsive to changes in neighborhood characteristics, which can in turn alter equilibrium rents through offsetting forces. Lower frictions increase out-migration incentives for residents currently in a given area j, putting downward pressure on rents in that neighborhood. At the same time, an increased incentive for households elsewhere to move to neighborhood j puts upward pressure on prices. Because these channels work in opposite directions, the net effect on equilibrium prices–and hence welfare—is difficult to sign a priori.

In Panel A of Appendix Table A.11, we report results from re-estimating the model under alternative calibrations of moving costs, reducing costs to 50 percent of the baseline value or eliminating them entirely. The resulting counterfactuals continue to show redistributive welfare effects, with income-and race-based disparities persisting across calibrations. While the aggregate impact of demolitions is negative under all calibrations, reductions in moving frictions yields substantive gains in welfare for many groups, particularly low-income Black households. Panel B suggests the mechanism: lower moving costs attenuate citywide rent growth—driven by smaller spillover rent increases in non-demolition tracts—because households can re-sort across neighborhoods more readily in response to demolition-induced shocks.

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

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

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

Allowing for Intensive Margin Housing Consumption Adjustment: Our baseline approach assumes that households consume a fixed unit of housing and does not explicitly model how households adjust the quantity of housing consumption within a location. A natural alternative is to allow households to adjust housing quantity within neighborhoods—for example, by assuming Cobb-Douglas preferences over a composite good and housing services, so that households choose both a neighborhood and an optimal level of housing consumption (e.g., Baum-Snow and Hartley, 2020; Davis et al., 2021). 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.

Under the Cobb–Douglas framework, the coefficient in the indirect utility function on the log housing price variable is equal to the housing expenditure share. As a result, we calibrate the rent sensitivity for each race and income group using the 2000 Census microdata rather than estimate it via our instrumental-variable strategy.[37] We continue to instrument for neighborhood demographic composition using the same approach as in our main analysis. Relative to our baseline model, this reduces reliance on the rent instrument but imposes stronger structure on within-neighborhood housing demand. Put differently, the choice between the baseline model and the Cobb–Douglas specification involves a tradeoff between placing more weight on identification through the IV strategy for rent versus imposing functional-form restrictions on housing consumption behavior.

To assess whether this additional structure is empirically warranted in our setting, we evaluate the Cobb-Douglas specification using the same out-of-sample validation exercise used in Figure 7. Intuitively, if allowing an intensive margin provides a better description of sorting behavior and

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

equilibrium outcomes, it should improve predictive performance when we estimate the model on one period and forecast neighborhood shares in another. In practice, we find that the Cobb-Douglas approach performs worse in our out-of-sample validation exercise for 1990 data, with the R^2 falling from 0.67 to 0.56 (Appendix Figure A.8).[38] We view this as strong evidence that our baseline model is better able to capture the economic mechanisms of how people are consuming and sorting within the city. A potential explanation for this weaker performance is the additional theoretical structure imposed by Cobb-Douglas preferences on within-neighborhood housing demand and the estimation of preference parameters. In particular, Cobb-Douglas restricts how housing consumption responds to rents and income, and it ties down rent sensitivity through a calibrated expenditure share. 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.

8 Evaluating Alternative Housing Policies

A natural next question is how additional housing policy responses influence the welfare consequences of public housing demolitions for both displaced public housing households and the broader set of Cook County residents. In this section, we use the estimated structural model to study the effects of demand- and supply-side interventions that might mitigate the disparate impacts of demolitions. On the demand side, we consider two targeted policies aimed at displaced public-housing residents. First, we increase housing voucher generosity, which changes the effective rent schedules displaced households face and expands the set of neighborhoods they can feasibly access. Second, we consider intensive housing counseling that reduces relocation frictions by lowering the moving costs associated with searching for and leasing a new unit. In terms of supply, we vary the scale of redevelopment in demolition-targeted areas, a broader intervention that affects displaced public housing residents and the broader population living in the city through equilibrium changes in rents and sorting. 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. For each policy, we present results separately for displaced public-housing residents and the broader population, highlighting how the same intervention can have sharply different welfare implications across these groups.

Welfare Effects Under Generous Housing Vouchers and Intensive Counseling: Welfare losses among displaced public housing residents are a key driver of the aggregate welfare consequences that we estimate. At the same time, the analysis in Table 4 shows that Section 8 vouchers improve these individuals' welfare by lowering the cost of housing in all neighborhoods and expanding choice. We consider two targeted demand-side interventions aimed at displaced residents: (i)

[38] 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.

increasing voucher generosity and (ii) providing intensive housing counseling that reduces moving costs. For voucher generosity, we increase the baseline maximum monthly voucher amount of $1,036 by 10% or 20%. In terms of counseling, we rely on Galiani, Murphy and Pantano (2015), who incorporate mobility counseling into a structural choice model by allowing it to reduce moving costs. They model counseling as an additive shifter in moving costs and identify it using the MTO experimental group that received mobility counseling; their estimates imply an effect of roughly one-tenth of baseline moving frictions, which motivates our calibration of intensive counseling as a 10 percent reduction in displaced residents' moving costs.[39] We assume that neither of these targeted policies affect the welfare of the broader set of Cook County residents.

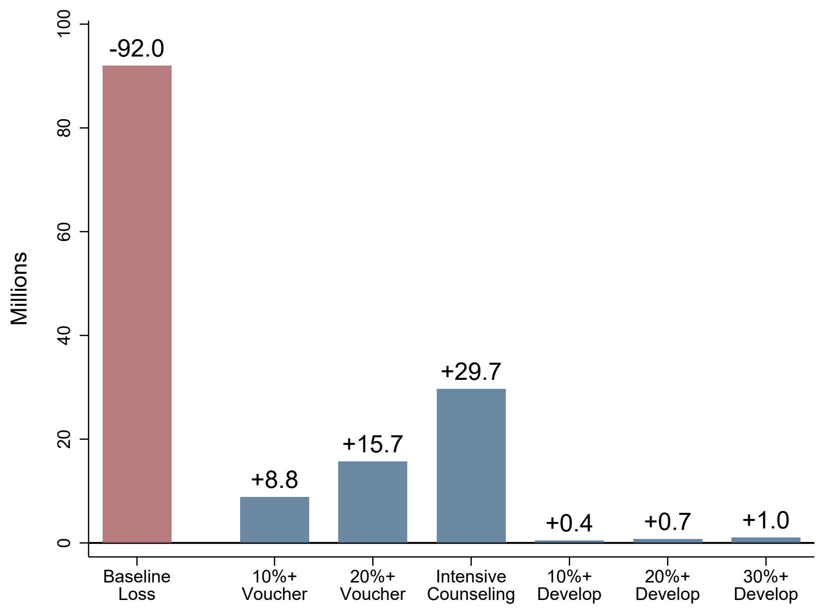
Figure 10 summarizes how these targeted demand-side policies affect the welfare of displaced public-housing residents. The first bar (maroon) on the left shows the magnitude of the baseline welfare loss from demolitions for displaced residents, which totals $92.0 million in rent-equivalent terms. Subsequent bars (navy) report the improvement relative to this baseline scenario under alternative voucher or counseling policies. Increasing voucher generosity by 10 percent raises displaced residents' welfare by $8.7 million relative to the baseline, while a 20 percent increase raises welfare by $15.6 million. The incremental gain from moving from a 10 percent to a 20 percent increase is $6.9 million, indicating diminishing returns. Consistent with this interpretation, the maximum voucher payment under the standard Section 8 payment schedule covers full rent in 41 percent of neighborhoods; this share rises to 62 percent under a 10 percent increase, but only 74 percent under a 20 percent increase. Even so, more generous vouchers do not eliminate the baseline losses for displaced residents: a 20 percent increase still leaves displaced residents with a net welfare loss of $76.3 million (= $92.0 − $15.7). By contrast, intensive counseling that reduces moving costs by 10 percent produces a larger attenuation, lowering the displaced-resident welfare loss to $62.3 million in rent-equivalent terms; this larger effect is consistent with the fact that baseline losses are driven primarily by relocation frictions.

Of course, increasing the generosity of housing vouchers would also require additional government expenditures. In our simulations, increasing voucher generosity by 10 percent raises voucher expenditures by $3.3 million compared to our baseline (which has total voucher expenditures of $143.7 million), while increasing generosity by 20 percent raises expenditures by $5.8 million relative to baseline. Thus, the rent-equivalent welfare of public-housing residents rises by more than expenditures on vouchers (which do not include the deadweight loss associated with raising these revenues). Intensive counseling would likewise require program resources. While the administrative costs of counseling in our setting are difficult to approximate, we use the MTO experiment as a benchmark to obtain an order-of-magnitude sense of program outlays. Goering et al. (1999) report that the net cost per family counseled at the Chicago MTO site was $1,683$ in 2017 dollars. Applying this benchmark to the set of displaced households in our analysis implies program outlays of roughly $23.5 million for an intensive counseling intervention of comparable

[39]In Galiani, Murphy and Pantano (2015), Table 3 reports estimates of the baseline moving-cost term and the counseling shifter that imply counseling offsets roughly one-tenth of baseline moving costs (approximately $0.68/6.60 \approx 0.10$.

scope.

Figure 10: Counterfactual Aggregate Welfare Impacts for Displaced Public-Housing Residents



Notes: This figure presents welfare results in millions of rent-equivalent dollars. The first bar on the left reproduces our main estimate of the welfare loss (indicated by maroon coloring) of public housing demolition for displaced households relative to a no-demolition counterfactual. All remaining bars show the increase in welfare associated with estimating the impact of demolition in alternative policy scenarios. Voucher policies increase voucher generosity (payment standard) for displaced households; an intensive housing counseling program reduces mobility costs by 10 percent; redevelopment policies increase construction in redevelopment neighborhoods.

Welfare Effects Under Additional Housing Redevelopment: Next, we consider a broader intervention on the supply side of the housing market: generating additional redevelopment in neighborhoods where public housing was demolished. Specifically, we consider a scenario where policymakers target areas with demolitions by creating additional market-rate housing units. We characterize redevelopment in terms of the share of total public housing that was destroyed. For example, redevelopment of 10 percent means that the government constructs additional market-rate housing in demolition neighborhoods that replaces 10 percent of the units that we observe as having been destroyed by 2010.

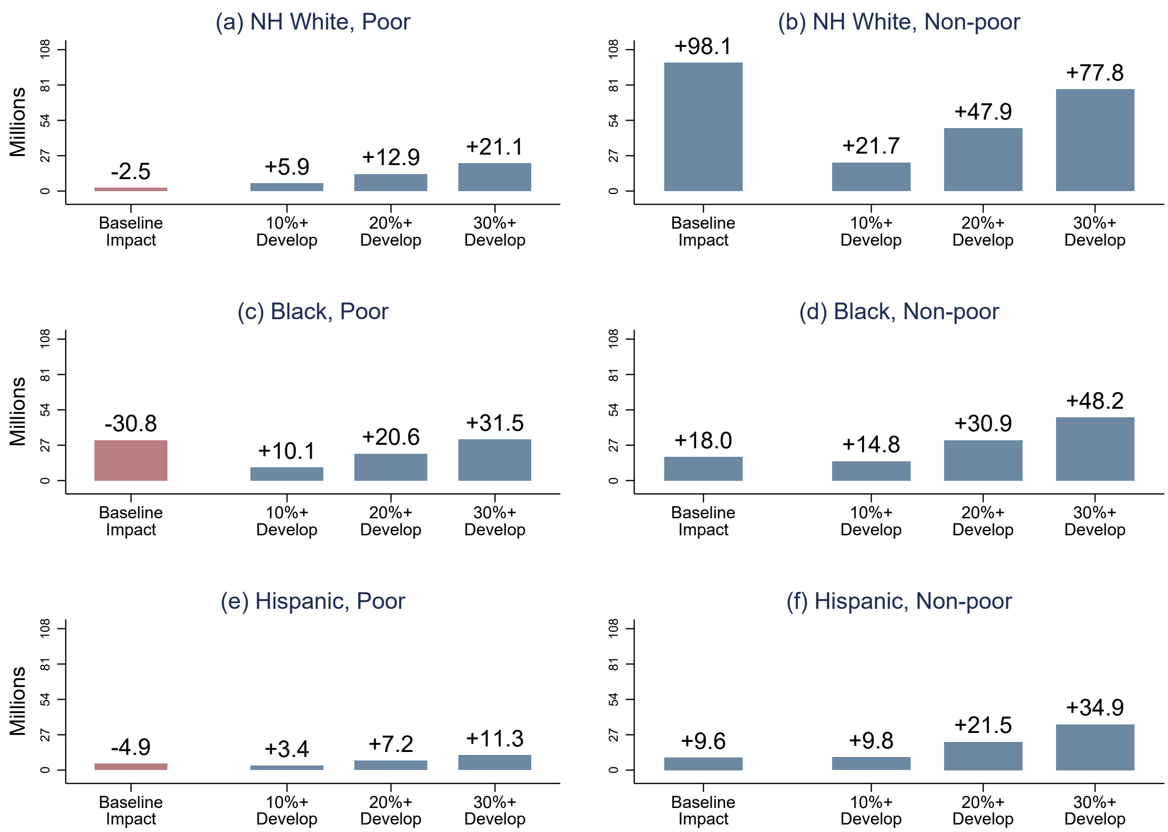
The motivation for this analysis stems from two considerations. First, as discussed in Section 2, redevelopment under the Plan for Transformation was uneven, with demolition often not being followed by new construction; consistent with this, many former public-housing sites remained only partially redeveloped years later, leaving substantial vacant and undeveloped land and making additional construction in these areas feasible (see Appendix Table A.1). Second, our earlier decomposition results show that demolitions affect the broader population through general-equilibrium changes in rents and neighborhood composition. As shown in Section 7.3, these effects arise because demolition removes a disamenity in targeted neighborhoods, raising rents directly in demolition tracts and indirectly through spillovers citywide. A natural implication is that expanding housing supply in demolition neighborhoods can mitigate rent pressure and reshape equilibrium sorting, with consequences for both displaced households and non-displaced residents.[40] We do not incorporate local government financing or construction costs, so the counterfactual should be interpreted as the equilibrium consequences of a more expansive federally sponsored HOPE VI-style redevelopment response.

We begin by assessing how redevelopment in demolition neighborhoods affects the welfare of displaced public-housing households. Despite being motivated by the original HOPE VI vision of rebuilding in demolition-targeted areas, the model implies that additional redevelopment has only modest benefits for displaced households. As shown in Figure 10, additional construction amounting to 10, 20, or 30 percent of the number of demolished units improves displaced households' aggregate welfare by just \$0.3, \$0.7, and \$1.0 million relative to the baseline demolition policy, leaving their overall losses essentially unchanged (given the baseline loss of \$92.0 million). These limited welfare improvements occur even though redevelopment materially unwinds the price response to demolitions. Appendix Figure A.9 shows that a 10 percent redevelopment expansion reduces the demolition-neighborhood rent impact from 15.9 percent to 8.3 percent, and a 20 percent expansion reduces it to 1.9 percent (with the 30 percent case nearly eliminating the rent effect in demolition neighborhoods). The intuition is that the primary source of displaced households' welfare loss is the disutility of forced relocation out of public-housing units, while the voucher-based rent schedule already insulates them from a portion of market-rent variation. As a result, place-based supply expansions in a limited set of demolition tracts do not substantially offset the bundled displacement costs borne by these households.

In sharp contrast to what we find for displaced public-housing households, we find that expanded redevelopment generates large welfare gains for the broader set of Cook County residents. Figure 11 reports impacts of alternative redevelopment policies for each of the race-by-income groups we consider. Across all groups, additional redevelopment produces sizable improvements relative to the baseline demolition policy, with the largest gains accruing to non-poor White households. Unlike displaced households—whose welfare is dominated by relocation costs and partly insulated by the voucher schedule—the unsubsidized population in Cook County is broadly exposed to equilibrium rents, so the price responses to redevelopment translate into large welfare changes in aggregate. Importantly, the effects for disadvantaged renter-heavy groups at sufficiently large scales of redevelopment can fully counter losses: for poor Black households, the baseline welfare loss is \$32.7 million, while redevelopment yields improvements of \$10.8 million, \$21.9 million, and \$33.2 million under 10, 20, and 30 percent redevelopment. These patterns reflect general-equilibrium adjustment: additional construction in demolition neighborhoods reduces equilibrium rent pressure both directly in demolition neighborhoods and indirectly by dampening spillover-driven rent increases elsewhere in the city. Because renter-heavy groups bear the

[40] In this scenario, we allow for this government redevelopment to crowd out private housing construction. The response of the latter is always summarized by the supply function in equation (5).

Figure 11: Counterfactual Aggregate Welfare Impacts for Cook County Residents Under Redevelopment



Notes: Each panel reports (i) the magnitude of the baseline welfare impact of demolitions (relative to a no-demolition counterfactual) and (ii) the *improvement relative to the baseline demolition policy* from increased redevelopment construction for six income-by-race groups. Maroon and navy coloring indicate negative and positive values, respectively. For instance, the left-most bar for Black, poor households shows that public housing demolition reduces welfare by -$32.7 million in rent-equivalent units. Redevelopment bars report policy improvements (the impact of demolitions in the policy counterfactual minus the baseline demolition scenario). When the navy improvement bar exceeds the maroon baseline bar, the policy more than offsets the baseline loss (i.e., the net effect becomes positive). Bar labels report exact values.

incidence of the price impacts of demolition, the effects of redevelopment disproportionately raise their welfare. Over the range we consider, the marginal gains from redevelopment are increasing, consistent with nonlinear responses of equilibrium rents and sorting to larger supply expansions once redevelopment is large enough to substantially unwind rent increases throughout the city.

Aggregate Welfare Effects of Alternative Policies: Our results indicate that housing voucher generosity, intensive housing counseling, and additional housing redevelopment—the policies most directly tied to a spatially targeted intervention such as public housing demolition—operate through distinct incidence channels. As noted in Section 7.1, summing the estimated baseline demolition impacts for displaced households and the six race-by-income groups yields an overall welfare loss of −\$4.6 million in rent equivalent terms. This net loss is driven by the large welfare decline for displaced public housing households (−\$92.0 million), which is only partly offset by gains among non-displaced residents.

The counterfactuals in this section show that targeted demand-side interventions can substantially mitigate the welfare losses experienced by displaced public-housing residents. Increasing voucher generosity expands the set of price-feasible neighborhoods and mitigates displaced residents' losses, but the incremental gains diminish as higher voucher payments increasingly cover rent in neighborhoods that are already close to being feasible. By contrast, counseling directly reduces relocation frictions, and because baseline losses for displaced residents are driven primarily by moving costs, this channel produces a larger attenuation. At the aggregate level, these targeted interventions matter because displaced households account for a dominant share of the baseline welfare loss: by partially offsetting displaced residents' losses while leaving broader-market outcomes essentially unchanged, both policies can reverse the sign of the overall welfare impact. For example, the baseline countywide welfare impact of demolition is a loss of −\$4.6 million, which becomes a gain of \$11.0 million when voucher generosity expands by 20 percent. Similarly, under an intensive housing counseling program that reduces moving costs by 10 percent, demolitions generate an overall welfare gain \$25.1 million. Importantly, however, even under these targeted demand-side policies, displaced residents remain worse off on net, underscoring that mitigating displacement costs is not the same as eliminating them.

Redevelopment has the opposite incidence. Additional construction in demolition neighborhoods does little to offset displaced households' losses (an improvement of only \$0.4–\$1.0 million across the redevelopment scenarios), but it generates large aggregate gains for the broader population through general-equilibrium changes in rents and sorting. As shown in Figure 11, scaling redevelopment construction by 10, 20, and 30 percent improves total welfare of the broader population by \$65.8, \$141.0, and \$224.7 million relative to the baseline demolition policy, yielding net totals of \$66.2, \$141.7, and \$225.7 million, respectively. These results highlight a central distinction: voucher generosity is the lever that most directly reduces demolition-driven welfare losses for displaced households, while redevelopment primarily matters because it reshapes equilibrium outcomes for the much larger non-displaced population.

9 Conclusion

This paper studies one of the largest place-based housing redevelopment initiatives in U.S. history: the HOPE VI program and the public housing demolitions it financed. These demolitions led to lasting changes in housing markets and the demographic composition of targeted neighborhoods (Tach and Emory, 2017; Staiger, Palloni and Voorheis, 2024). Focusing on Chicago—a landmark HOPE VI site—we combine administrative records that track displaced public housing residents and their relocation destinations with Census/ACS data on neighborhood characteristics, prices and citywide sorting. Using these data, we estimate a structural model of neighborhood choice to quantify welfare impacts, with particular attention to the relocation experiences of displaced households and to the distributional incidence across racial and income groups.

Our main finding is that demolitions had sharply disparate impacts. Displaced public-housing households experience large welfare losses driven by the disutility of forced relocation, and although vouchers expand choice sets, the welfare gains from expanded choice are too small to offset relocation costs. For the broader Cook County population, the average effects differ sharply by income and tenure status. Higher-income groups gain on net—with the largest gains accruing to non-poor White households—while poorer, renter-heavy groups experience losses, especially poor Black and Hispanic households. These unequal effects of demolitions in the non-public housing population arise from two important forces. First, while all households benefit from the destruction of public housing as a neighborhood attribute, demolition-induced changes in neighborhood composition and subsequent re-sorting generate heterogeneous welfare impacts across groups. Second, equilibrium rent increases (both directly in demolition tracts and via spillovers) shift welfare from renters to homeowners; because homeownership rates differ sharply by race and income, price appreciation amplifies disparities in incidence.

Overall, our results highlight a central challenge for place-based revitalization policies that are intended to benefit lower-income households: the same interventions that raise neighborhood amenities can also raise equilibrium rents and induce re-sorting, shifting benefits toward households that are better positioned to capture them. In our setting, the welfare incidence hinges on two margins that are often discussed (Glaeser and Gottlieb, 2008; Neumark and Simpson, 2015) but rarely quantified jointly: displacement costs borne by the households forced to move, and equilibrium price changes that transfer welfare from renters to homeowners and alter who can afford to live in improved neighborhoods. As a result, neighborhood "improvements" are not synonymous with welfare gains for disadvantaged residents—especially when relocation is costly and when housing supply responses are limited.

Our structural approach contributes to studies of place-based policies by placing displaced households and the broader market in the same welfare accounting framework: we combine direct evidence on relocation behavior with an equilibrium model that captures rent spillovers and endogenous sorting. The broader lesson is that evaluating neighborhood redevelopment requires simultaneously measuring (a) what happens to displaced residents, (b) how prices respond citywide, and (c) how those price changes reshape neighborhood composition. These forces are likely to matter in other cities and programs where redevelopment changes amenities at scale, triggers displacement, and interacts with sharp differences in tenure, price sensitivity, and preferences over neighborhood composition.

Finally, our counterfactuals clarify that alternative housing policies have distinct impacts on the welfare consequences of urban renewal programs such as public housing demolitions. More generous vouchers primarily benefit displaced households by expanding feasible choices, although they only partially offset relocation costs. Empirically-relevant reductions in moving costs via counseling for displaced residents generate similar conclusions. Redevelopment in demolition areas, by contrast, has little effect on displaced households and instead matters through general-equilibrium changes in rents and sorting that affect the much larger non-displaced population. In our simulations, vouchers or housing counseling can reduce displaced losses and can make aggregate (countywide) welfare positive at larger increases in generosity, whereas expanded redevelopment produces large gains for the broader population and—at sufficiently large scales—can fully offset baseline losses for renter-heavy groups. This framing also changes how to interpret "insufficient redevelopment" historically: HOPE VI sites with limited redevelopment—as documented in major cities such as Atlanta and Washington, D.C. (Vale, Shamsuddin and Kelly, 2018)—likely realized smaller general-equilibrium gains from demolition, limiting the extent to which renter-heavy groups shared in those gains.

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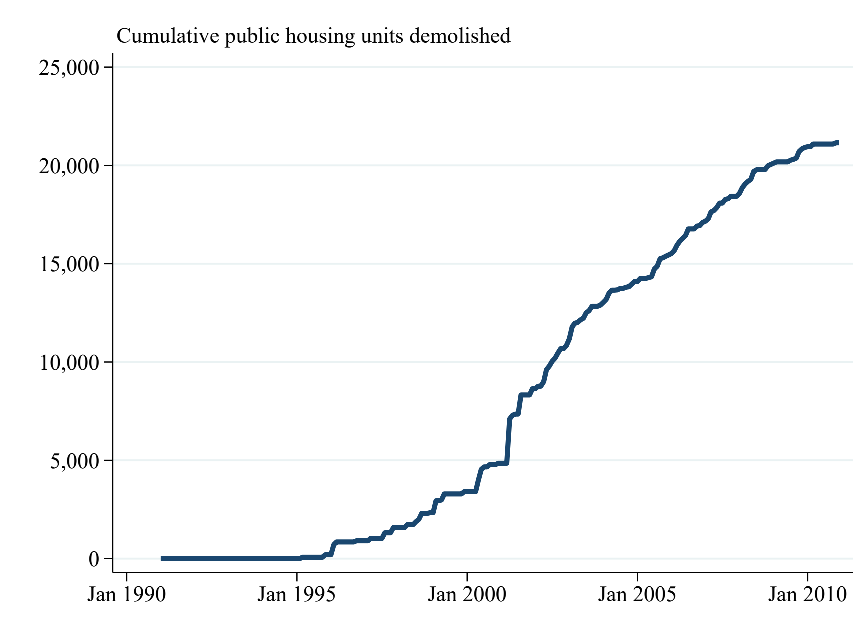
Online Appendices

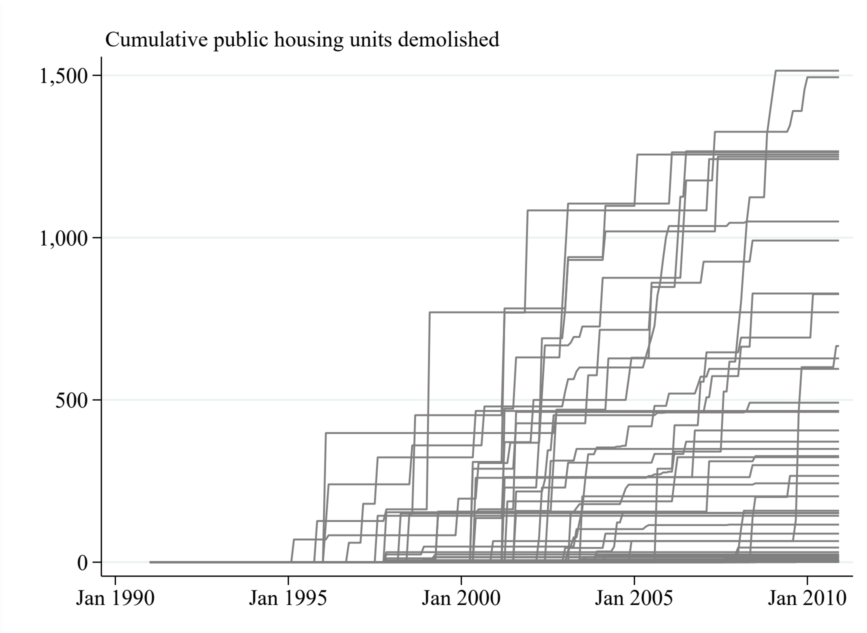
A Appendix Figures and Tables

Appendix Figure A.1: Time Series of Public Housing Demolitions in Chicago

(a) Chicago

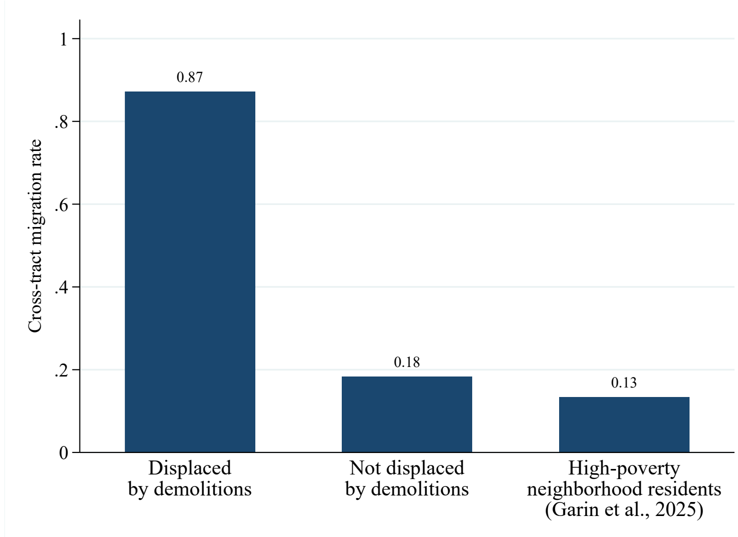


(b) By Tract



Notes: Panel A displays the cumulative number of public housing units that were demolished in Chicago between 1995 and 2010. Panel B displays results separately for each of the 59 census tracts that experienced a demolition.
Source: Authors' calculations using data from the Chicago Housing Authority.

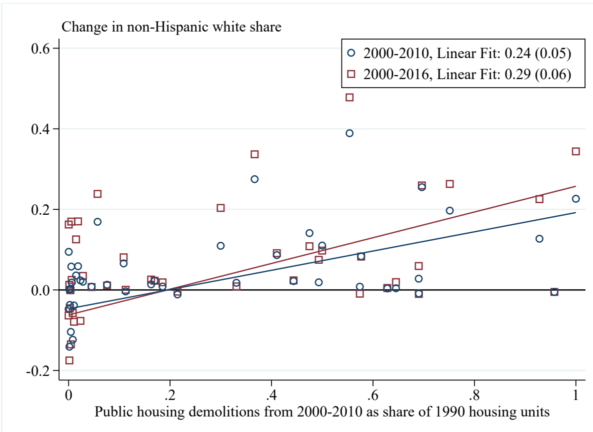
Appendix Figure A.2: Migration Rates, Public Housing Sample



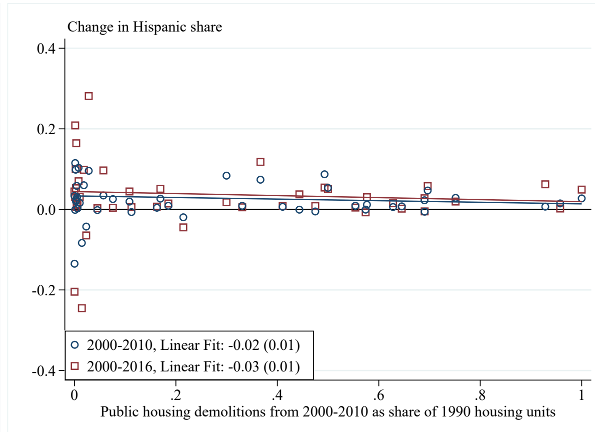
Notes: The first two bars of this figure report the share of households that moved across census tracts after public housing demolitions, separately for households that were displaced from their housing unit by demolitions and households that were not. For this sample, locations are measured one year after demolition for 84% of the sample and two years after for the remaining 16%. The final bar reports a comparable cross-tract migration rate for people who start out in neighborhoods where the poverty rate exceeds 30% from Garin et al. (2025).

Appendix Figure A.3: Supplementary Neighborhood Outcomes and Public Housing Demolitions, 2000–2010 and 2000–2016

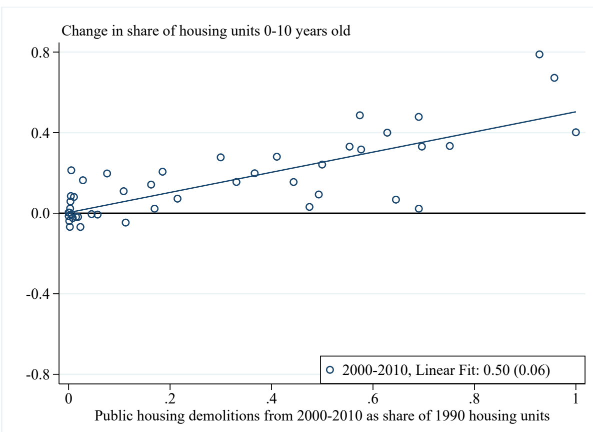
(a) Non-Hispanic White Population Share



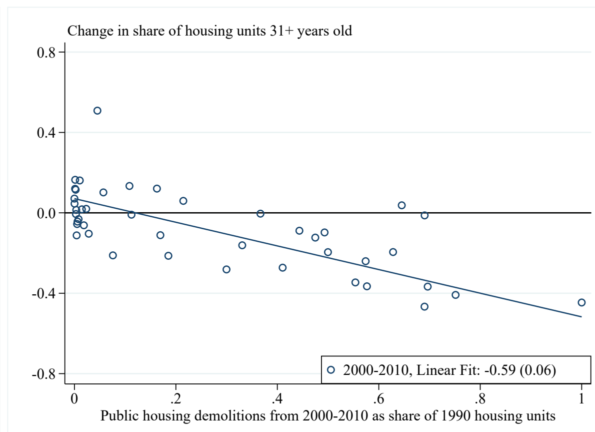
(b) Hispanic Population Share



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



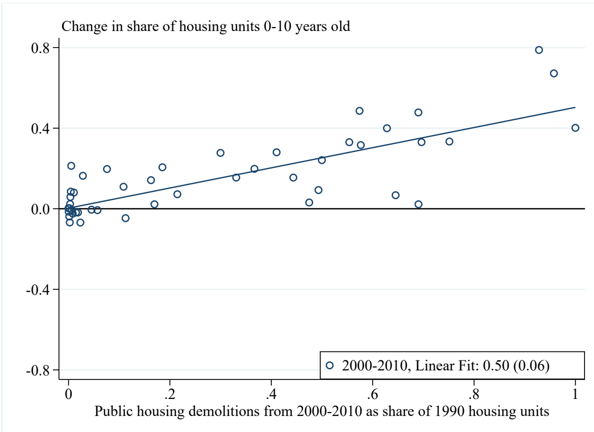
(d) Share of Housing Units 31+ Years Old



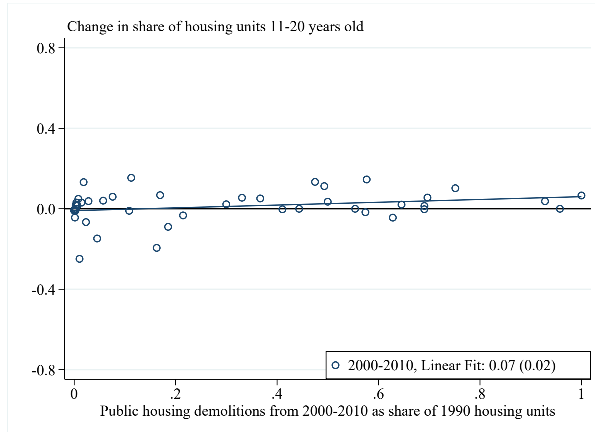
Notes: This figure plots the change in neighborhood characteristics against the cumulative number of public housing units demolished from 2000–2010 as a share of the number of occupied housing units in 1990. Each dot represents the average change in the indicated dependent variable for a given discrete value of the extent of public housing demolition. We winsorize the public housing demolition share variable from above at 1 for 3 tracts. Panels (c) and (d) do not contain 2000–2016 changes because the available ACS data for 2016 do not allow us to construct consistent housing-unit age bins.
Source: Authors' calculations using data from the Chicago Housing Authority and U.S. Census Bureau.

Appendix Figure A.4: Changes in Housing Stock and Public Housing Demolitions, 2000–2010

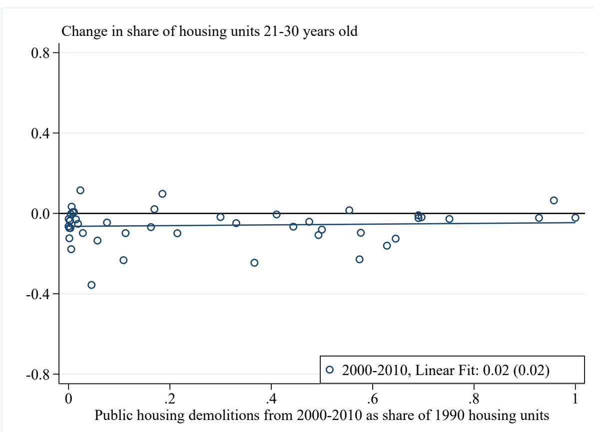
(a) Share of Housing Units 0–10 Years Old



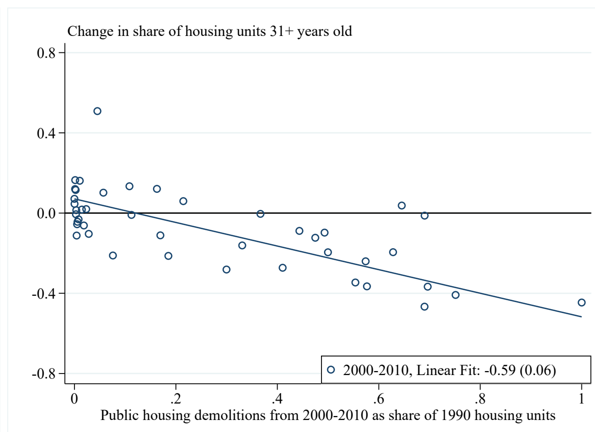
(b) Share of Housing Units 11-20 Years Old



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



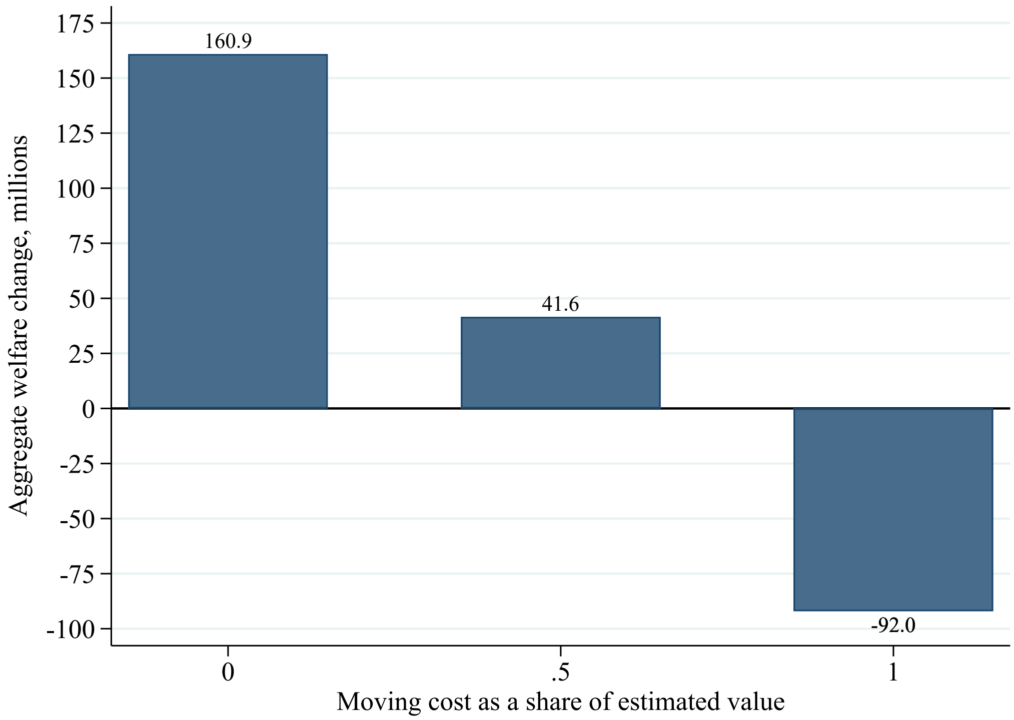
(d) Share of Housing Units 31+ Years Old



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.

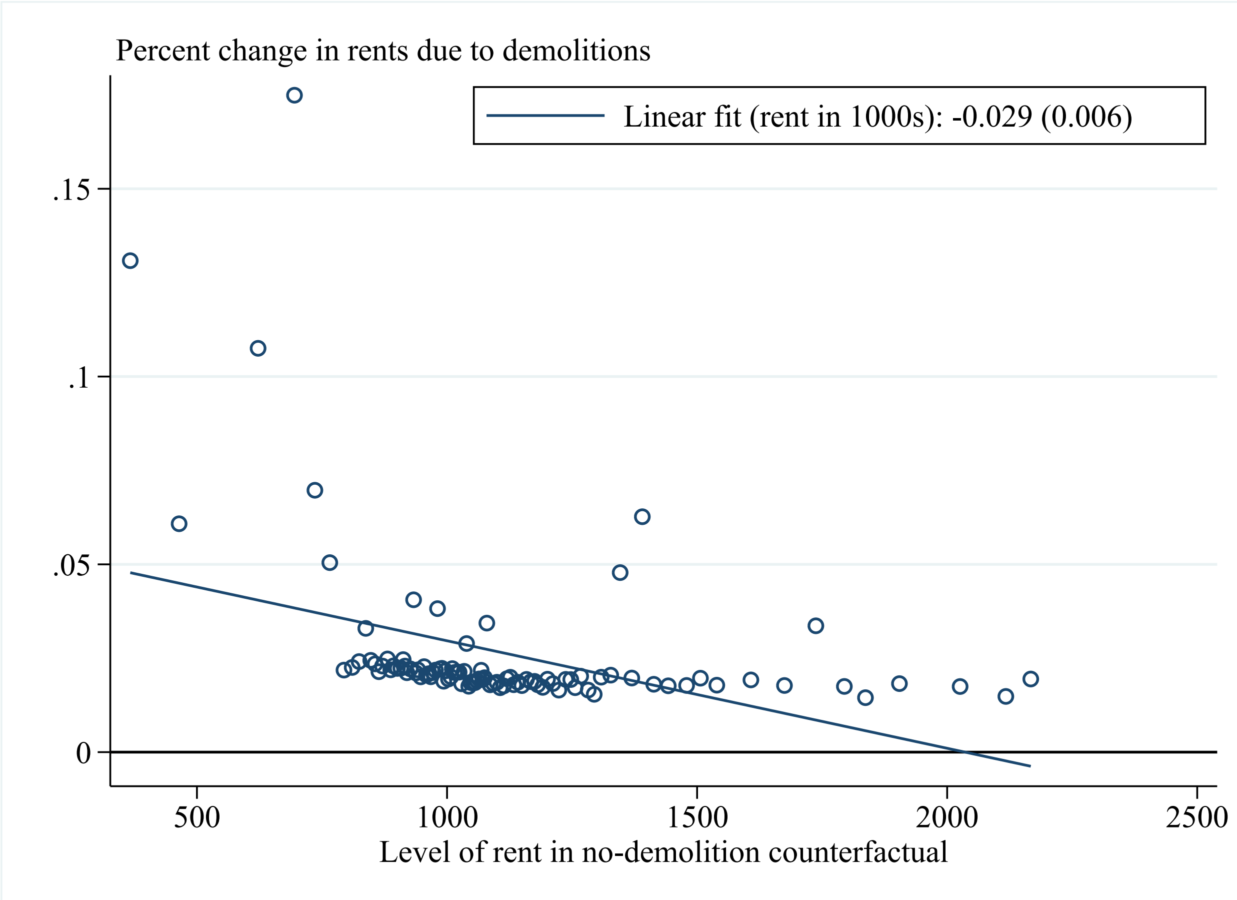
Appendix Figure A.5: Sensitivity of Welfare Effects of Public Housing Demolitions for Public Housing Residents to Estimated Moving Cost



Notes: This figure reports the aggregate rent equivalent change in welfare, measured in millions of dollars, of public housing residents from the actual state of the world in 2010 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. Our baseline estimates are shown in the right-most bar. The other bars show results when the unit and neighborhood moving costs are scaled down to 50% of their estimated value or zero.

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

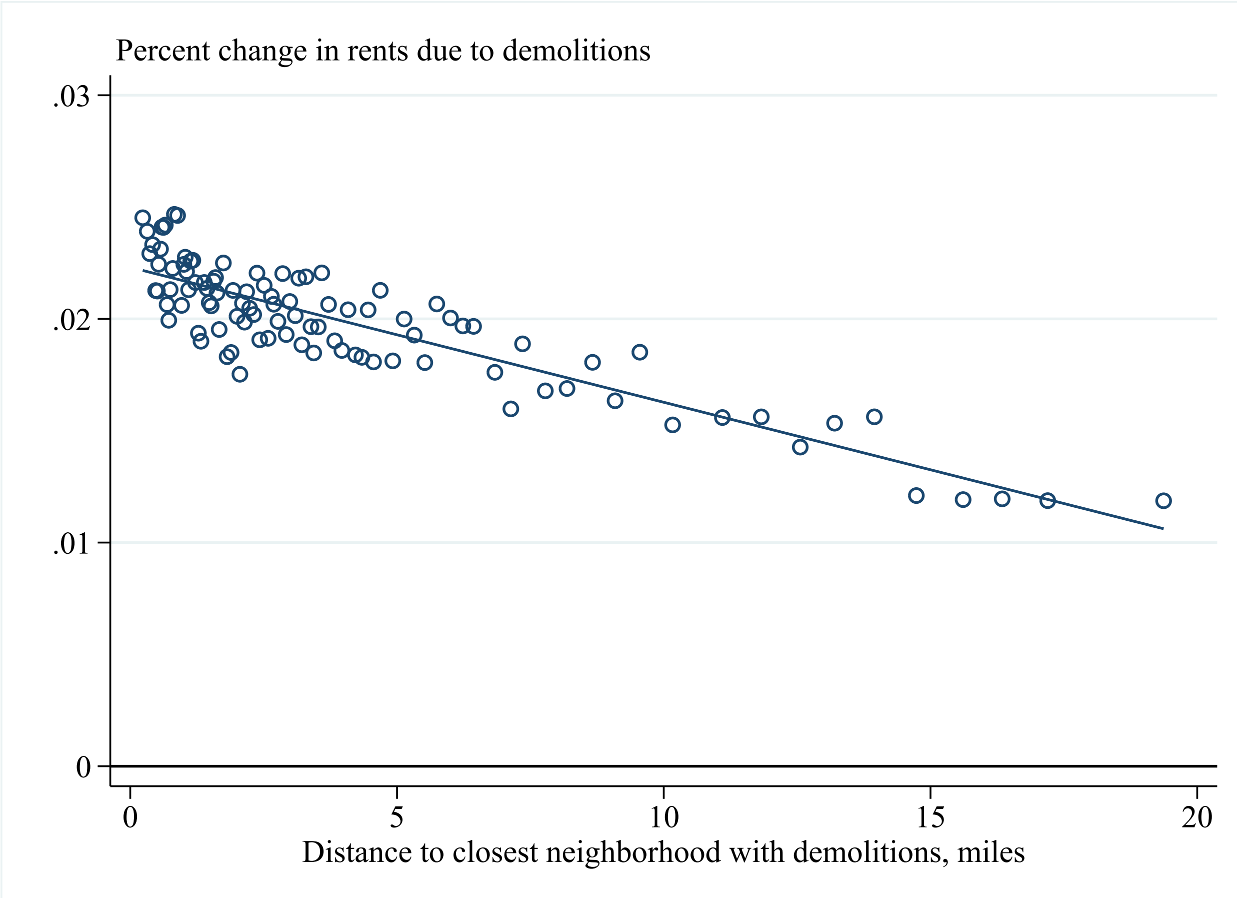
Appendix Figure A.6: Tract-Level Rent Changes Due to Public Housing Demolitions Relative to Rents in Absence of Demolitions



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.7: Rents Increased by More in Non-Public-Housing Neighborhoods That Were Closer to Public Housing Demolitions

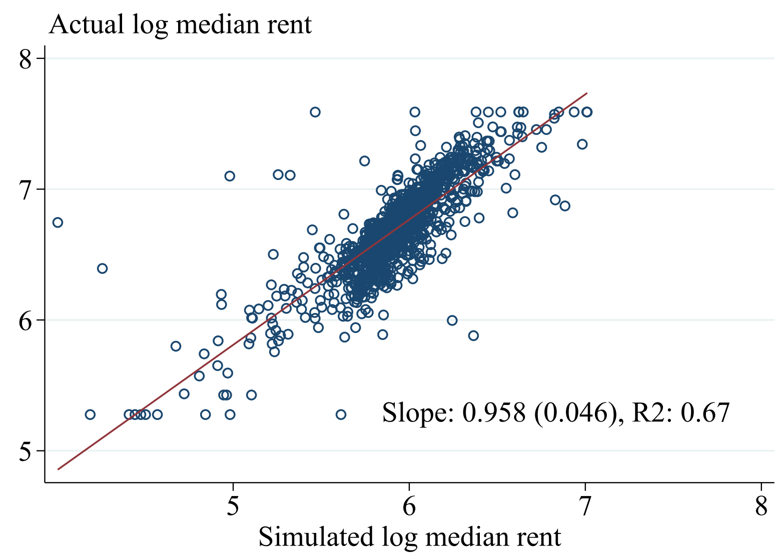


Notes: Figure displays the percent change in median rents due to public housing demolitions for tracts that did not have public housing against the distance to the closest tract with demolitions. We construct the dependent variable using estimates from the model and a comparison of differences between 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.8: Comparing Out-of-Sample Fit of Baseline Model and Cobb-Douglas Model Using Rent Data

(a) Baseline Model



(b) Cobb-Douglas Model

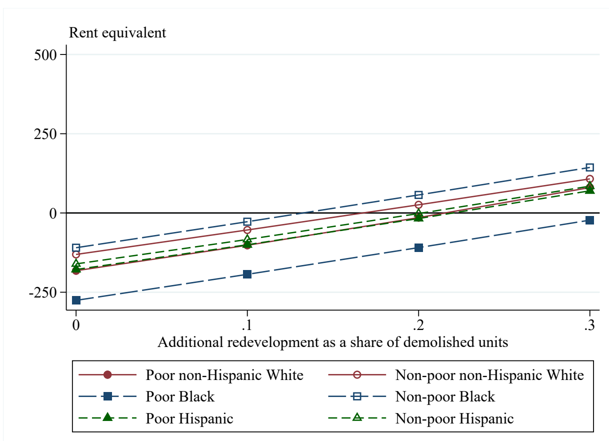
Notes: Panel A repeats the plot from Figure 7 of actual log rents in 1990 in census tracts against log rents that are simulated by an out-of-sample procedure. In particular, we construct simulated rents for 1990 using the coefficients and tract fixed effects estimated using 2000–2010 data, exogenous 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 (6)–(8). 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.

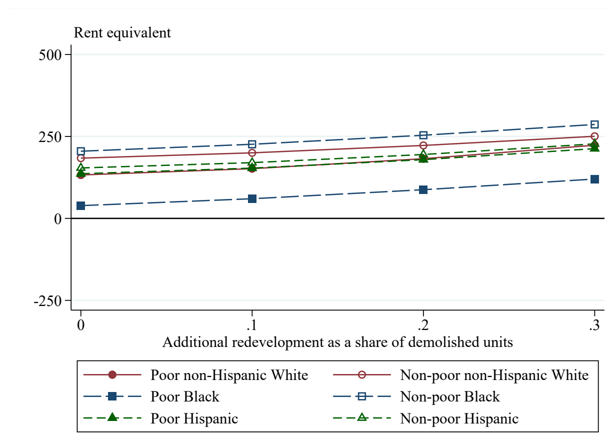
Appendix Figure A.9: Consequences of Public Housing Demolitions on Housing Prices Under Additional Redevelopment of Public Housing

(a) Housing Prices

(b) Average Welfare Consequences for Renters



(c) Average Welfare Consequences for Owners



Notes: Panel A displays the percent change in rents due to demolitions under counterfactual scenarios in which the indicated share of demolished public housing units in each neighborhood are rebuilt by the government. Panel B shows the rent equivalent welfare effect of public housing demolitions for renters and Panel C 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 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: Baseline Sample Characteristics, Public Housing Sample

	Not displaced by demolitions (1)	Displaced by demolitions (2)
Female	0.86	0.88
	(0.34)	(0.32)
Black	0.99	0.99
	(0.09)	(0.09)
Age	32.54	32.17
	(11.56)	(10.33)
Household size	2.17	2.13
	(1.44)	(1.37)
At least one child	0.51	0.51
	(0.50)	(0.50)
At least one child under age 5	0.06	0.06
	(0.25)	(0.24)
Earnings	5102.33	5219.71
	(1444.63)	(1438.85)
Observations	2459	507

Notes: Table reports means and standard deviations (in parentheses) of baseline characteristics of the public housing analysis sample. Column 1 contains results for individuals who were not displaced by the early demolitions. Column 2 contains results for individuals who were displaced.
Source: Authors' calculations using data from the Chicago Housing Authority.

Appendix Table A.3: OLS Estimates of Neighborhood Preference Parameters

	Preference parameters for indicated group		
	Non-Hispanic White	Black	Hispanic
	(1)	(2)	(3)
Panel A: Poor Households			
Log median rent	-0.0452***	-0.0280***	-0.0362***
	(0.00572)	(0.00464)	(0.00751)
Black population share	-0.0857***	0.196***	0.239***
	(0.0162)	(0.0171)	(0.0225)
Hispanic population share	-0.0491***	0.0260**	0.232***
	(0.0141)	(0.0122)	(0.0207)
Log median household income	0.00247	-0.00512	-0.0243***
	(0.00390)	(0.00466)	(0.00678)
PH units as a share of housing stock	-0.00988	-0.0100	-0.0309
	(0.0136)	(0.0184)	(0.0238)
Panel B: Non-poor Households			
Log median rent	-0.000432	-0.000786	-0.0168***
	(0.00180)	(0.00267)	(0.00576)
Black population share	-0.0966***	0.161***	0.151***
	(0.00720)	(0.0149)	(0.0181)
Hispanic population share	-0.103***	0.0536***	0.148***
	(0.00405)	(0.00699)	(0.0166)
Log median household income	0.00632***	-0.00140	-0.0174***
	(0.00167)	(0.00273)	(0.00544)
PH units as a share of housing stock	-0.0110*	-0.0402***	-0.0429**
	(0.00648)	(0.0113)	(0.0176)
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.4: Instrumental Variable Estimates of Neighborhood Preference Parameters, Poor Non-Hispanic White and Black Households, Robustness to Controls and Instruments

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	Add 1990-2000 controls (4)	Drop < 1 mile Cabrini-Green (5)	IV: Education, No income (6)	IV: Add slope & bedrock (7)
Poor Non-Hispanic White Households							
Log median rent	-0.446***	-0.445***	-0.449***	-0.376***	-0.453***	-0.479***	-0.422***
	(0.0717)	(0.0727)	(0.0717)	(0.0445)	(0.0748)	(0.0619)	(0.0627)
Black population share	-0.134	-0.0929	-0.130	-0.0811	-0.0968	-0.123	-0.122
	(0.0817)	(0.0799)	(0.0830)	(0.0570)	(0.0800)	(0.0827)	(0.0749)
Hispanic population share	-0.0623	-0.0350	-0.0620	-0.0620*	-0.0791*	-0.0727	-0.0473
	(0.0454)	(0.0483)	(0.0455)	(0.0349)	(0.0473)	(0.0535)	(0.0425)
Log median household income	0.0841***	0.0889***	0.0847***	0.0960***	0.0764***	0.0945***	0.0824***
	(0.0225)	(0.0256)	(0.0226)	(0.0165)	(0.0220)	(0.0240)	(0.0216)
PH units as a share of housing stock	-0.439***	-0.427***	-0.433***	-0.324***	-0.455***	-0.475***	-0.411***
	(0.111)	(0.110)	(0.110)	(0.0757)	(0.120)	(0.110)	(0.104)
Poor Black Households							
Log median rent	-0.0602**	-0.0551***	-0.0602**	-0.0593***	-0.110***	-0.309***	-0.136***
	(0.0244)	(0.0169)	(0.0244)	(0.0197)	(0.0267)	(0.0419)	(0.0255)
Black population share	0.342***	0.229***	0.342***	0.342***	0.353***	0.332***	0.292***
	(0.0223)	(0.0176)	(0.0225)	(0.0223)	(0.0250)	(0.0547)	(0.0295)
Hispanic population share	0.0916***	0.0791***	0.0916***	0.0874***	0.0720***	-0.0143	0.0717***
	(0.0177)	(0.0139)	(0.0177)	(0.0182)	(0.0196)	(0.0366)	(0.0196)
Log median household income	0.0203***	0.0141***	0.0203***	0.0229***	0.0265***	0.0795***	0.0320***
	(0.00654)	(0.00500)	(0.00655)	(0.00742)	(0.00697)	(0.0168)	(0.00825)
PH units as a share of housing stock	-0.0591*	-0.0614**	-0.0585*	-0.0525*	-0.0994**	-0.332***	-0.135***
	(0.0352)	(0.0251)	(0.0348)	(0.0301)	(0.0427)	(0.0766)	(0.0426)

Notes: This table presents estimates of preference parameters for the broader Cook County sample based on 2SLS estimation of equation (10). 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 treated as endogenous. Column 1 reports results from our baseline specification (also reported in Table 2). 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. Column 4 adds interactions between year fixed effects and changes from 1990–2000 in median log rents, median log household income, and share of adults with a college degree. Column 5 drops tracts that are within 1 mile of the Cabrini-Green Homes. Column 6 replaces the income-based shift-share instrumental variables with education-based ones (grade 12, 1–2 years of college, and 4+ years of college). Column 7 adds to our baseline instrumental variables interactions between year fixed effects and the mean slope of land in a tract and the mean bedrock depth. 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.5: Instrumental Variable Estimates of Neighborhood Preference Parameters, Poor Hispanic and Other Race/Ethnicity Households, Robustness to Controls and Instruments

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	Add 1990-2000 controls (4)	Drop < 1 mile Cabrini-Green (5)	IV: Education, No income (6)	IV: Add slope & bedrock (7)
Poor Hispanic Households							
Log median rent	-0.272***	-0.143***	-0.265***	-0.256***	-0.353***	-0.409***	-0.312***
	(0.0520)	(0.0331)	(0.0510)	(0.0363)	(0.0628)	(0.0568)	(0.0514)
Black population share	0.287***	0.236***	0.289***	0.281***	0.317***	0.315***	0.263***
	(0.0563)	(0.0359)	(0.0556)	(0.0479)	(0.0636)	(0.0729)	(0.0622)
Hispanic population share	0.306***	0.326***	0.307***	0.248***	0.267***	0.280***	0.316***
	(0.0367)	(0.0271)	(0.0362)	(0.0335)	(0.0429)	(0.0515)	(0.0398)
Log median household income	0.0291*	0.00738	0.0279*	0.0427***	0.0338*	0.0737***	0.0373**
	(0.0155)	(0.0112)	(0.0152)	(0.0135)	(0.0184)	(0.0227)	(0.0176)
PH units as a share of housing stock	-0.289***	-0.157***	-0.282***	-0.233***	-0.371***	-0.437***	-0.328***
	(0.0801)	(0.0543)	(0.0771)	(0.0621)	(0.105)	(0.102)	(0.0862)
Poor Other Race/Ethnicity Households							
Log median rent	-0.309***	-0.0112	-0.310***	-0.224***	-0.253***	0.303***	-0.0512
	(0.0730)	(0.0324)	(0.0726)	(0.0533)	(0.0675)	(0.0554)	(0.0400)
Black population share	-0.489***	-0.0613**	-0.485***	-0.441***	-0.460***	-0.290***	-0.342***
	(0.0877)	(0.0257)	(0.0880)	(0.0682)	(0.0785)	(0.0535)	(0.0439)
Hispanic population share	-0.217***	0.0780***	-0.216***	-0.231***	-0.211***	0.0489	-0.112***
	(0.0510)	(0.0288)	(0.0509)	(0.0488)	(0.0488)	(0.0506)	(0.0343)
Log median household income	0.0168	-0.00742	0.0172	0.0203	-0.000530	-0.0933***	-0.0197
	(0.0210)	(0.00901)	(0.0210)	(0.0186)	(0.0181)	(0.0206)	(0.0121)
PH units as a share of housing stock	-0.463***	-0.0496	-0.455***	-0.336***	-0.437***	0.193**	-0.195***
	(0.107)	(0.0408)	(0.107)	(0.0744)	(0.101)	(0.0857)	(0.0579)

Notes: See notes to Appendix Table A.4.

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 Non-Hispanic White and Black Households, Robustness to Controls and Instruments

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	Add 1990-2000 controls (4)	Drop < 1 mile Cabrini-Green (5)	IV: Education, No income (6)	IV: Add slope & bedrock (7)
Non-Poor Non-Hispanic White Households							
Log median rent	-0.0646***	-0.0652***	-0.0671***	-0.0529***	-0.0424***	-0.0356***	-0.0315***
	(0.0133)	(0.0122)	(0.0135)	(0.00935)	(0.0105)	(0.00712)	(0.00787)
Black population share	-0.114***	-0.0751***	-0.113***	-0.0947***	-0.107***	-0.119***	-0.0989***
	(0.0147)	(0.0126)	(0.0151)	(0.0106)	(0.0116)	(0.0104)	(0.00899)
Hispanic population share	-0.137***	-0.104***	-0.137***	-0.126***	-0.131***	-0.136***	-0.128***
	(0.00886)	(0.00851)	(0.00903)	(0.00820)	(0.00757)	(0.00743)	(0.00613)
Log median household income	0.0176***	0.0171***	0.0181***	0.0202***	0.0134***	0.00768***	0.0129***
	(0.00380)	(0.00406)	(0.00392)	(0.00317)	(0.00269)	(0.00250)	(0.00237)
PH units as a share of housing stock	-0.0811***	-0.0659***	-0.0808***	-0.0627***	-0.0611***	-0.0507***	-0.0461***
	(0.0184)	(0.0175)	(0.0189)	(0.0126)	(0.0140)	(0.0113)	(0.0112)
Non-Poor Black Households							
Log median rent	-0.0655***	0.0124	-0.0645***	-0.0685***	-0.0824***	-0.0851***	-0.0441***
	(0.0158)	(0.0128)	(0.0155)	(0.0122)	(0.0179)	(0.0139)	(0.0125)
Black population share	0.174***	0.205***	0.176***	0.175***	0.184***	0.188***	0.182***
	(0.0222)	(0.0160)	(0.0222)	(0.0212)	(0.0236)	(0.0231)	(0.0192)
Hispanic population share	0.0638***	0.112***	0.0643***	0.0390***	0.0515***	0.0654***	0.0791***
	(0.0104)	(0.0102)	(0.0103)	(0.0105)	(0.0116)	(0.0130)	(0.00932)
Log median household income	0.0107**	-0.00319	0.0106**	0.0183***	0.0112**	0.0197***	0.00845**
	(0.00445)	(0.00393)	(0.00441)	(0.00467)	(0.00486)	(0.00542)	(0.00382)
PH units as a share of housing stock	-0.114***	-0.0145	-0.110***	-0.0999***	-0.132***	-0.135***	-0.0893***
	(0.0257)	(0.0184)	(0.0251)	(0.0208)	(0.0308)	(0.0274)	(0.0210)

Notes: See notes to Appendix Table A.4.

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, Non-Poor Hispanic and Other Race/Ethnicity Households, Robustness to Controls and Instruments

	Baseline (1)	Add spatial controls (2)	Add murder rate (3)	Add 1990-2000 controls (4)	Drop < 1 mile Cabrini-Green (5)	IV: Education, No income (6)	IV: Add slope & bedrock (7)
Non-Poor Hispanic Households							
Log median rent	-0.271***	-0.0964***	-0.264***	-0.250***	-0.323***	-0.216***	-0.241***
	(0.0492)	(0.0268)	(0.0480)	(0.0347)	(0.0576)	(0.0342)	(0.0407)
Black population share	0.0975*	0.152***	0.101*	0.0953**	0.124**	0.154***	0.115**
	(0.0569)	(0.0280)	(0.0561)	(0.0473)	(0.0613)	(0.0422)	(0.0501)
Hispanic population share	0.151***	0.225***	0.152***	0.0884***	0.119***	0.201***	0.185***
	(0.0332)	(0.0228)	(0.0326)	(0.0303)	(0.0378)	(0.0317)	(0.0309)
Log median household income	0.0237	-0.00204	0.0227	0.0378***	0.0239	0.0258**	0.0213
	(0.0152)	(0.00895)	(0.0148)	(0.0130)	(0.0169)	(0.0130)	(0.0139)
PH units as a share of housing stock	-0.316***	-0.118***	-0.307***	-0.252***	-0.378***	-0.256***	-0.281***
	(0.0781)	(0.0429)	(0.0751)	(0.0590)	(0.0964)	(0.0598)	(0.0688)
Non-Poor Other Race/Ethnicity Households							
Log median rent	0.222***	0.314***	0.217***	0.207***	0.316***	0.550***	0.402***
	(0.0475)	(0.0549)	(0.0470)	(0.0348)	(0.0593)	(0.0748)	(0.0635)
Black population share	-0.242***	-0.0612	-0.244***	-0.241***	-0.248***	-0.197**	-0.155**
	(0.0423)	(0.0550)	(0.0419)	(0.0359)	(0.0527)	(0.0891)	(0.0706)
Hispanic population share	-0.237***	-0.0565	-0.238***	-0.197***	-0.193***	-0.105	-0.185***
	(0.0359)	(0.0392)	(0.0357)	(0.0339)	(0.0429)	(0.0678)	(0.0481)
Log median household income	-0.0381***	-0.0541***	-0.0374***	-0.0489***	-0.0472***	-0.108***	-0.0655***
	(0.0138)	(0.0191)	(0.0136)	(0.0125)	(0.0171)	(0.0295)	(0.0223)
PH units as a share of housing stock	0.131*	0.286***	0.125*	0.0862	0.232**	0.488***	0.320***
	(0.0682)	(0.0825)	(0.0663)	(0.0548)	(0.0930)	(0.132)	(0.105)

Notes: See notes to Appendix Table A.4.

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

Appendix Table A.8: Instrumental Variable Estimates of Neighborhood Preference Parameters, Poor Households, Robustness to Calibrated Moving Cost

	Moving cost as a percent of baseline value				
	100%	90%	75%	50%	0%
	(1)	(2)	(3)	(4)	(5)
Poor Non-Hispanic White Households					
Log median rent	-0.446***	-0.452***	-0.461***	-0.471***	-0.473***
	(0.0717)	(0.0576)	(0.0547)	(0.0539)	(0.0536)
Black population share	-0.134	-0.127*	-0.126	-0.130*	-0.132*
	(0.0817)	(0.0771)	(0.0766)	(0.0791)	(0.0797)
Hispanic population share	-0.0623	-0.0845**	-0.0944**	-0.0984**	-0.0998**
	(0.0454)	(0.0424)	(0.0440)	(0.0468)	(0.0476)
Log median household income	0.0841***	0.0855***	0.0862***	0.0875***	0.0875***
	(0.0225)	(0.0221)	(0.0217)	(0.0219)	(0.0220)
PH units as a share of housing stock	-0.439***	-0.451***	-0.464***	-0.475***	-0.477***
	(0.111)	(0.105)	(0.102)	(0.103)	(0.103)
Poor Black Households					
Log median rent	-0.0602**	-0.186***	-0.211***	-0.218***	-0.220***
	(0.0244)	(0.0269)	(0.0291)	(0.0295)	(0.0297)
Black population share	0.342***	0.301***	0.269***	0.258***	0.257***
	(0.0223)	(0.0373)	(0.0413)	(0.0432)	(0.0437)
Hispanic population share	0.0916***	0.0789***	0.0961***	0.102***	0.103***
	(0.0177)	(0.0221)	(0.0248)	(0.0265)	(0.0270)
Log median household income	0.0203***	0.0364***	0.0345***	0.0325***	0.0322***
	(0.00654)	(0.00983)	(0.0107)	(0.0110)	(0.0111)
PH units as a share of housing stock	-0.0591*	-0.189***	-0.207***	-0.209***	-0.209***
	(0.0352)	(0.0501)	(0.0538)	(0.0548)	(0.0551)
Poor Hispanic Households					
Log median rent	-0.272***	-0.269***	-0.236***	-0.219***	-0.217***
	(0.0520)	(0.0404)	(0.0373)	(0.0362)	(0.0362)
Black population share	0.287***	0.247***	0.227***	0.237***	0.242***
	(0.0563)	(0.0537)	(0.0472)	(0.0449)	(0.0447)
Hispanic population share	0.306***	0.339***	0.364***	0.369***	0.368***
	(0.0367)	(0.0339)	(0.0322)	(0.0318)	(0.0318)
Log median household income	0.0291*	0.0196	0.00719	0.00208	0.00113
	(0.0155)	(0.0151)	(0.0136)	(0.0131)	(0.0130)
PH units as a share of housing stock	-0.289***	-0.283***	-0.246***	-0.232***	-0.230***
	(0.0801)	(0.0730)	(0.0668)	(0.0651)	(0.0652)

Notes: This table presents estimates of preference parameters for the broader Cook County sample based on 2SLS estimation of equation (10). Column 1 reports results from our baseline specification (also reported in Table 2). In columns 2–5, we construct the mean indirect utility, δ^k_{jt}, when scaling down the baseline moving cost to 90%, 75%, 50%, and 0% of its baseline value (see footnote 20). 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.9: Instrumental Variable Estimates of Neighborhood Preference Parameters, Non-Poor Households, Robustness to Calibrated Moving Cost

	Moving cost as a percent of baseline value				
	100%	90%	75%	50%	0%
	(1)	(2)	(3)	(4)	(5)
Non-Poor Non-Hispanic White Households					
Log median rent	-0.0646***	-0.0567***	-0.0811***	-0.0945***	-0.0965***
	(0.0133)	(0.00959)	(0.0115)	(0.0126)	(0.0127)
Black population share	-0.114***	-0.0990***	-0.104***	-0.115***	-0.118***
	(0.0147)	(0.0120)	(0.0155)	(0.0180)	(0.0185)
Hispanic population share	-0.137***	-0.143***	-0.154***	-0.159***	-0.159***
	(0.00886)	(0.00779)	(0.00991)	(0.0115)	(0.0118)
Log median household income	0.0176***	0.0210***	0.0274***	0.0302***	0.0306***
	(0.00380)	(0.00344)	(0.00446)	(0.00509)	(0.00520)
PH units as a share of housing stock	-0.0811***	-0.0722***	-0.0974***	-0.110***	-0.112***
	(0.0184)	(0.0147)	(0.0186)	(0.0208)	(0.0211)
Non-Poor Black Households					
Log median rent	-0.0655***	-0.0320***	-0.0582***	-0.0780***	-0.0834***
	(0.0158)	(0.00962)	(0.0116)	(0.0135)	(0.0141)
Black population share	0.174***	0.211***	0.220***	0.221***	0.222***
	(0.0222)	(0.0186)	(0.0212)	(0.0241)	(0.0250)
Hispanic population share	0.0638***	0.0964***	0.114***	0.122***	0.123***
	(0.0104)	(0.00916)	(0.0110)	(0.0126)	(0.0132)
Log median household income	0.0107**	0.00907***	0.0162***	0.0211***	0.0225***
	(0.00445)	(0.00341)	(0.00418)	(0.00488)	(0.00509)
PH units as a share of housing stock	-0.114***	-0.0781***	-0.105***	-0.126***	-0.132***
	(0.0257)	(0.0185)	(0.0223)	(0.0256)	(0.0266)
Non-Poor Hispanic Households					
Log median rent	-0.271***	-0.152***	-0.119***	-0.107***	-0.106***
	(0.0492)	(0.0269)	(0.0254)	(0.0258)	(0.0260)
Black population share	0.0975*	0.116***	0.119***	0.134***	0.140***
	(0.0569)	(0.0342)	(0.0281)	(0.0265)	(0.0264)
Hispanic population share	0.151***	0.230***	0.262***	0.271***	0.271***
	(0.0332)	(0.0231)	(0.0218)	(0.0216)	(0.0217)
Log median household income	0.0237	0.00105	-0.00633	-0.00821	-0.00842
	(0.0152)	(0.00990)	(0.00901)	(0.00888)	(0.00890)
PH units as a share of housing stock	-0.316***	-0.179***	-0.138***	-0.126***	-0.125***
	(0.0781)	(0.0467)	(0.0418)	(0.0414)	(0.0417)

Notes: See notes to Appendix Table A.8.

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

Appendix Table A.10: 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	0.159	0.020	0.026
Share of total rent increase	0.29	0.71	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.11: The Aggregate Welfare and Rent Effects of Public Housing Demolitions, Robustness to Calibrated Moving Cost for Broader Population

	Moving cost as a percent of baseline value		
	100%	50%	0%
	(1)	(2)	(3)
Panel A: Aggregate welfare change, millions			
Public housing residents	-92.0	-93.6	-93.5
Cook County residents			
Non-Hispanic White, Poor	-2.5	-1.0	0.0
Non-Hispanic White, Non-poor	98.1	83.4	85.5
Black, Poor	-30.8	-18.3	-16.4
Black, Non-poor	18.0	16.1	17.6
Hispanic, Poor	-4.9	-3.8	-3.1
Hispanic, Non-poor	9.6	12.1	13.7
Total	-4.6	-5.2	3.8
Panel B: Average percent change in rents			
Tracts with demolitions	0.159	0.167	0.167
Tracts without demolitions	0.020	0.017	0.016
All tracts	0.026	0.024	0.023

Notes: Panel A reports the aggregate rent equivalent change in welfare, measured in millions of dollars, 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. Panel B reports the average percent change in rents due to demolitions. Column 1 reports results from our baseline value of the calibrated moving cost for the broader set of Cook County residents. Columns 2 and 3 report results when the moving cost is 50% of the baseline value or 0.
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.[41]

[41] 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.

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

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

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

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

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

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