Who Gains When Neighborhoods Rebuild? Redevelopment, Sorting, and Housing Affordability*

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Abstract

How does housing supply affect affordability and welfare when new, housing replaces the existing stock? We document that redevelopment upgrades housing quality without adding units and raises neighborhood housing prices and average income. We then develop a dynamic spatial model with forward-looking landlords renovating or redeveloping a depreciating stock, income sorting across neighborhoods and housing qualities, and endogenous neighborhood amenities. We estimate the redevelopment elasticity using quasi-experimental variation from an anti-redevelopment policy in Chicago. We find that redevelopment lowers rents for high-quality housing yet raises rents for low-quality housing. Restricting redevelopment in gentrifying neighborhoods leaves all but the poorest households worse off and shifts redevelopment to untreated low-income neighborhoods. Endogenous amenities shape the local and city-wide effects of redevelopment and of policies seeking to reduce housing costs.

Keywords: Housing redevelopment, income sorting, housing affordability, gentrification, anti-redevelopment policy.

JEL Codes: R21, R31, R38

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1 Introduction

Understanding housing supply is central to the affordability challenge. In dense, built-up cities, new supply comes largely from redevelopment, which demolishes an existing building and replaces it with a new one (McMillen and O'Sullivan, 2013; Baum-Snow and Han, 2024). Many cities restrict redevelopment to protect low-income residents,[1] yet its effects on housing affordability and welfare are not well understood. New high-quality units can lower rents citywide through vacancy chains (Mast, 2023; Nathanson, 2025); but, in contrast with greenfield development, the loss of old, lower-quality units can raise local rents and displace incumbent residents. How does redevelopment affect affordability across neighborhoods and for households of varying incomes? How should cities design housing policy when new supply replaces the existing stock?

We develop a dynamic spatial general equilibrium model to answer these questions. In the model, forward-looking landlords make endogenous redevelopment decisions, and redevelopment in turn attracts high-income households into the neighborhood. Absent redevelopment, housing quality depreciates and filters down to low-income households over time (Brueckner and Rosenthal, 2009). We estimate the model using data from Chicago, where redevelopment is prevalent and pervades policy discourse (Institute for Housing Studies, 2021). We discipline the model with the causal effect of a recent anti-redevelopment policy, as well as the salient differences in characteristics between old and new housing.

Our analysis yields three findings. First, redevelopment lowers rents for high-quality housing and raises them for low-quality housing, both locally and city-wide, so it raises the welfare of high-income households and lowers that of low-income ones. The "trickle-down" effect, whereby new high-quality construction lowers rents for lower-quality units (Nathanson, 2025), is limited: migrant high-income households occupy the new units, while the low-quality units are lost to redevelopment. Second, policies that restrict redevelopment can lower the city's average rent while eroding land values and reducing welfare for all but the poorest households. Falling rent coincides with lower welfare because housing quality deteriorates, so standard measures of rent burden, such as the average rent or the rent-to-income ratio, do not necessarily track welfare. Third, endogenous amenities, which reinforce income sorting once the quality distribution shifts, are central to understanding the distributional effects of redevelopment and of spatially targeted anti-redevelopment policies.

[1] For example, Chicago taxes demolition permits and bans redevelopment from reducing a parcel's unit count in selected neighborhoods; San Francisco requires conditional use authorization to demolish, merge, or convert residential units, restricts the conversion of rental buildings to condominiums, and mandates one-for-one replacement of demolished rent-controlled units; and Seattle conditions demolition permits on relocation assistance for displaced low-income tenants.

The empirical analysis establishes three facts about redevelopment using permit and assessment records for Chicago. First, redevelopment is the main source of new housing supply. About three-fifth of construction permits are associated with a demolition permit, and this share rises to about four-fifth where construction is most active. Across block groups, a one-percentage-point increase in the construction rate comes with a 0.80-percentage-point increase in the redevelopment rate. Second, redevelopment upgrades quality rather than increases housing units. It concentrates in neighborhoods with low incomes, high rents, and old housing, and at the parcel level the new buildings contain fewer, larger, and higher-quality units than the buildings they replace.[2] Third, redevelopment causes neighborhood gentrification. We instrument the redevelopment rate with the interaction of a shift-share variable of housing demand shock and the initial share of old housing, which governs how strongly a neighborhood redevelops when demand rises. We find that a one-standard-deviation increase in the redevelopment rate raises decadal growth in housing values and incomes by about 0.8 standard deviations each, with no increase in the number of housing units or in population.

Motivated by these empirical facts, we build a general equilibrium model that incorporates endogenous redevelopment choices and income sorting. The model builds on the assignment framework: indivisible housing units differ by quality, and rents at each quality level adjust to match higher income households to higher quality units (Määttänen and Terviö, 2014; Landvoigt, Piazzesi and Schneider, 2015). We extend this framework in three ways. First, we incorporate heterogeneous neighborhoods and endogenous neighborhood amenities that respond to average income (Guerrieri, Hartley and Hurst, 2013). Second, we incorporate redevelopment and renovation decisions by forward-looking landlords, so the housing quality distribution responds endogenously to the returns of quality upgrading.[3] Third, we make the model dynamic. Housing quality depreciates over time, generating a source of low-quality housing – a process known as filtering (Sweeney, 1974; Brueckner and Rosenthal, 2009). While the standard assignment framework takes the distributions of household income and housing quality as given, our model endogenizes both at the neighborhood level and over time. Therefore, the model is well suited to assess the local and city-wide effects of redevelopment, as well as policies that restrict or promote it.

In our model, redevelopment and income sorting are mediated by an important equilibrium object: the rent elasticity of quality. High-income households, all else equal, sort into neighborhoods where high-quality housing is relatively cheap; equivalently, where this elasticity is low. On the other hand, landlords are more likely to redevelop where high-quality

[2] The fall in units per parcel reflects the replacement of two- to four-unit buildings by single-family homes, which is common in Chicago and is the stated motivation for the anti-redevelopment policies we study.

[3] In the model, redevelopment tears down the structure and rebuilds it at a new quality and unit count; renovation raises the quality of the existing structure and leaves its unit count unchanged.

housing commands a higher rent premium; where this elasticity is high. In equilibrium, an exogenous increase in high-quality housing supply reduces the rent elasticity of quality, which makes the neighborhood more attractive to high income households.

We estimate the model using rich data on property assessments, transaction records, building permits, rental listings, and neighborhood demographics for Chicago. We first recover each housing unit's quality, the depreciation rate, and the renovation premium from a hedonic regression of market rents on building characteristics. Consistent with the model, we allow the empirical rent-to-quality relationship to vary across neighborhoods. We calibrate construction costs using redevelopment rates and the quality and housing unit density of new buildings relative to those they replaced. This means our model can reproduce the fact that new buildings tend to be of higher quality and of lower density. We choose income-specific amenities to match income distributions across neighborhoods.

Most importantly, we discipline the redevelopment elasticity σ_c, a parameter that governs how strongly landlords' redevelopment decisions respond to their returns, using quasi-experimental evidence. Between 2021 and 2024, Chicago implemented a set of anti-redevelopment policies in the 606 Trail and Pilsen neighborhoods: a surcharge on demolition permits of at least $\$15,000$ and an anti-deconversion ordinance barring redevelopment projects from reducing a parcel's housing unit count. The policy has been controversial, defended for preserving unsubsidized affordable housing and criticized for stalling investment in aging neighborhoods.[4] We apply two difference-in-differences designs, one comparing parcels just inside the policy boundary to those just outside, and one comparing treated block groups to matched controls elsewhere in the city, to estimate the policy's causal effect on redevelopment. We find that the policy cuts demolitions by more than half, and choose σ_c so that a model counterpart of the policy matches this treatment effect.

We use the estimated model to assess an expanded version of the anti-redevelopment policy in 606 Trail and Pilsen. This counterfactual has direct policy relevance: Chicago has enacted the Northwest Side Preservation Ordinance, which extends the anti-deconversion rule to more neighborhoods and raises the surcharge to \$60,000 per demolition. We impose an expansion of this form for fifty years on neighborhoods with below-median income and above-median redevelopment rates. Where it applies, the policy works as intended. Redevelopment and gentrification fall sharply, the average rent falls as the housing stock deteriorates, and

[4] Nitkin (2022) reports competing views on the ordinances. Steven Vance, an urban planner, argues that "this ordinance worked exactly the way it was written to work" and should be expanded "to make sure we do not lose the type of housing that contains the bulk of the unsubsidized affordable housing in Chicago, which are the two-, three- and four-unit buildings." Paul Colgan, of the Homebuilders Association of Greater Chicago, counters that "[b]uildings deteriorate, and have to be either fixed up or replaced ... how do you improve the neighborhood if you don't invest in it?" and that "[t]here's been a chilling effect on investment in the neighborhood."

higher-income residents leave. However, the policy relocates redevelopment to untreated neighborhoods and causes gentrification in untreated neighborhoods; the largest spillovers occur in untreated neighborhoods that most resemble the treated ones. Citywide, rent falls except for top-quality housing, and the average rent falls by 2.3 percent, yet the policy costs households $1.63 billion in present-value welfare. While the lowest income households benefit substantially from the policy, all households making above $35,000 a year are made worse off because they consume lower quality housing. The dilapidation of housing also reduces land values and, as a consequence of the out-migration of high-income households, reduces neighborhood amenities.

Given the adverse consequences of the anti-redevelopment policy Chicago adopts, we use the model to evaluate two alternative policies that could improve housing affordability. A redevelopment subsidy in above-median-income neighborhoods generates the largest aggregate welfare gain by lowering the rent of medium and high-quality housing and improving the amenities of the treated neighborhoods. However, the policy is regressive because low-quality housing is lost to redevelopment, it fails to meaningfully expand housing unit supply, and it fails to deliver "trickle-down" affordability gains. In contrast, a rent subsidy in below-median-income neighborhoods is progressive among renters. However, a majority of the subsidy is capitalized into rents because new development is only marginally responsive. Through the lens of the model, welfare-improving housing policies should promote the sorting of high-income households into high-income neighborhoods because households do not internalize how their location choices affect neighborhood amenities. It should also support low-income households by promoting greater affordability in low-income neighborhoods.

Related Literature. We contribute to the literature on housing filtering. Sweeney (1974) and Brueckner (1980a) model housing as a depreciating asset that filters down the income distribution as it ages, with redevelopment timing determined by the option to rebuild (Brueckner, 1980b); Rosenthal (2008) and Brueckner and Rosenthal (2009) show that neighborhood income falls as the stock ages, and Rosenthal (2014) estimates a sizable filtering rate at the individual building level. We extend this literature by jointly modeling landlords' forward-looking redevelopment choices and heterogeneity in household housing demand, which together produce a dynamic feedback between quality upgrading via redevelopment and filtering.

Our model builds on assignment models of the housing market, in which vertically differentiated units are matched to households of different incomes (Määttänen and Terviö, 2014; Landvoigt, Piazzesi and Schneider, 2015).[5] Recent work extends it to ask how income-specific

[5]More broadly, the assignment framework has been used to study matching in the labor market (Costinot and Vogel, 2010), neighborhood choice (Davis and Dingel, 2020; Couture et al., 2024), and school choice (Epple and Romano, 2003), among other topics.

credit conditions (Nikolakoudis, 2024), mortgage interest rates (Hacamo, 2024), and housing supply and demand policies (Nathanson, 2025; Abramson and Landvoigt, 2025) affect housing costs across the distribution. We differ by endogenizing supply through forward-looking landlords, by making the framework dynamic — the relevant horizon given the durability and irreversibility of housing investment (Baum-Snow and Duranton, 2025) — and by embedding it in a quantitative spatial model with heterogeneous neighborhoods. This allows us to match the between-neighborhood variation in redevelopment we document and evaluate neighborhood-targeted policies.

Our work is related to the literature on how changes in housing supply affect local affordability, income and demographic composition, and neighborhood amenities. This work spans housing revitalization programs (Rossi-Hansberg, Sarte and Owens, 2010; Agarwal, Deng, Fan, Gao, Li and Ma, 2026), low-income housing development (Baum-Snow and Marion, 2009; Diamond and McQuade, 2019), market-rate building development (Pennington, 2021; Asquith, Mast and Reed, 2023; Mast, 2023; Kennedy and Wheeler, 2023; Mense, 2025), and public housing demolition (Almagro, Chyn and Stuart, 2026; Blanco, 2023). We provide a general equilibrium framework that jointly evaluates the local and city-wide effects of any change, in any neighborhood, in the quantity and quality distribution of housing. Peng (2023) and Rollet (2025) both formulate dynamic redevelopment models to study how floorspace responds to zoning reforms in New York City. We complement these two papers by incorporating the quality margin and indivisible housing units. In the New York City context, Lebret et al. (2026) use a partial equilibrium model to show that redevelopment policies raise rents by increasing housing quality. We complement this work by providing a general equilibrium framework whereby the distribution of housing costs responds endogenously to redevelopment policy.

Lastly, we build on the literature on quantitative spatial models (Ahlfeldt, Redding, Sturm and Wolf, 2015; Monte, Redding and Rossi-Hansberg, 2018; Hsieh and Moretti, 2019), including those with dynamics (Kleinman, Liu and Redding, 2023; Greaney, Parkhomenko and Van Nieuwerburgh, 2025). This class of models treats housing as divisible and clears the market by equating neighborhood-level floorspace demand and supply, abstracting from the granular choice over heterogeneous housing units that is standard in discrete-choice approaches to housing (Bayer, Ferreira and McMillan, 2007). We incorporate heterogeneous housing units into a quantitative spatial model and show that the housing quality distribution is itself an important determinant of income sorting across space.

Outline. Section 2 describes the data, Section 3 the empirical facts, Section 4 the model and its analytical results, Section 5 the estimation, and Section 6 the three policies. Section 7 concludes, and the appendices collect the proofs, the institutional detail on the Chicago policies, the parcel-matching procedure, and supporting results.

2 Data

Our analysis of the city of Chicago draws on five main data sources.

Property assessments and transactions. Annual assessment records from 2000 to 2023 are obtained from the Cook County Assessor's Office. Each record is identified by a unique parcel identification number (PIN) and contains building characteristics including build year, total floorspace, number of units, and quality attributes. The data cover single- and small multi-family buildings of up to six units; large apartment buildings are not included. Transaction deeds from the Cook County Recorder of Deeds record the universe of parcel transactions over the same period. We restrict the sample to arm's-length transactions—excluding sales below \$10,000, special deed types such as quit-claim or executor deeds, and duplicate records—and match transactions to assessment records by PIN.

Building permits. The City of Chicago Data Portal publishes all building permits issued from 2006 to 2023, recording the permit type, issue date, address, work description, and estimated cost. We restrict to residential permits and use the work description to classify each as a demolition, construction, or renovation permit, employing ChatGPT to identify permits that involve an entire housing unit rather than partial alterations such as roof repairs or electrical work.[6]

Rental listings. Residential rental listings come from RentHub, a nationwide dataset built by weekly scraping of more than 100 publicly accessible listing websites. We obtain the records in Chicago covering 2014–2024 at monthly frequency. Each listing records the posted rent, address, geographic coordinates, and building characteristics. We geocode all listings and match them to assessment records by address, deferring to the assessment records when building characteristics conflict. We drop listings with rents or floorspace below the 1st or above the 99th percentile, and collapse the data to the unit-year level by averaging posted rents within each year.

Block-group characteristics. Block-group-level employment by two-digit NAICS industry comes from the Residential Area Characteristics (RAC) files of the LODES dataset

[6]The prompt instructs: "Can you determine whether this description of a housing permit involves an entire housing unit? Adding or demolishing one or more floors of a building would count as a permit on an entire unit, such as a second-floor addition to an existing building. Some permits may involve work on only part of a unit, such as the roof or electrical system. Please return 1 if so and 0 if not."

(2005–2019), and demographic and income tabulations from the American Community Survey (ACS) (2009–2019). We also use the LODES origin–destination files for 2002, which record the number of workers commuting between each pair of census blocks, aggregated to block groups.

Neighborhood amenities. We draw from two data sources to measure neighborhoods amenities. We obtain the number of service establishments from the Data Axle Historical Business Database, which records the name, address, geographic coordinates, industry classification, employment, and sales of U.S. establishments at annual frequency. We compute the number of establishments by block group and year in sectors 44–45 (retail trade), 71 (arts, entertainment, and recreation), 72 (accommodation and food services), and 81 (other services). We obtain reported crime from the City of Chicago Data Portal, which publishes every incident recorded in the Chicago Police Department's system since 2001, and count incidents by block group and year over the same period.

Parcel definitions and permit matching. We measure redevelopment by matching building permits—how we detect redevelopment events—to land parcels, our unit of observation throughout the paper. Because parcels are divided, combined, and re-surveyed over time, often in response to development itself, we fix parcel boundaries at their definitions in the 2005 assessment files for the entire analysis period. By this definition, there are approximately 603,000 parcels within the Chicago municipality.

We match each building permit to a parcel in several steps. First, we match each permit to the assessment record with the same address and use the record that is closest in time to the permit date—the most recent record at or before a demolition or renovation permit, the earliest at or after a new-construction permit—and place that record's coordinates on the 2005 parcel map. Second, we match permits whose addresses do not match with any assessment record using the municipality's building footprints map, obtained from the City of Chicago Data Portal. Together, we successfully assign 98.5% of permits to a parcel.[7] Finally, because a 2005 lot may since have been divided or combined, we assign each new-construction permit to every 2005 lot overlapping the 2024 lot it was built on: a permit on an assembled lot attaches to several 2005 lots, and several permits on a subdivided lot attach to one. Appendix A.1 describes the full matching procedure in detail.

The geography of building activities. Building activity is spread unevenly across Chicago. Appendix Figure F.1 maps the construction, demolition, and redevelopment rates across

[7]The unmatched remainder consists primarily of renovation permits at addresses appearing in neither an assessment nor the official building shapefile. These tend to be permits for airports, hospitals and schools.

Chicago block groups. We define each rate as the share of 2005 parcels receiving a permit of that type in a year, averaged over 2009–2019, and count a parcel as redeveloped only when its demolition permit is matched to a new-construction permit on the same parcel. Construction is concentrated on the north side and in a few southern neighborhoods close to the Loop. Demolition is more dispersed: it is high in the same northern neighborhoods, but also across a large part of the south side where almost no construction takes place. Redevelopment concentrates where construction and demolition rates are both high.[8] In the next section, we analyze this spatial heterogeneity and study how redevelopment affects neighborhoods.

3 Empirical Facts on Housing Redevelopment

We use building permit and assessment data from Chicago to answer three questions about redevelopment. To what extent is redevelopment important for housing supply? Where does it occur and how does it reshape the housing supply? How does it affect neighborhood-level outcomes? In answering these questions, we focus on one-to-six unit buildings covered by our assessment data, which account for about three quarters all new construction permits and over half of the entire stock of housing units in the municipality.[9]

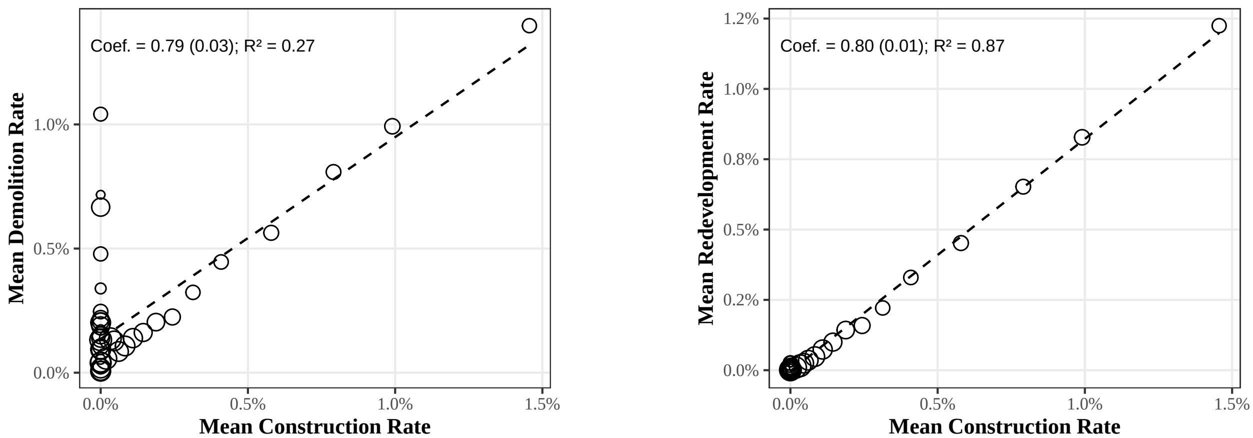
3.1 Redevelopment as a Primary Source of Housing Supply

We start by examining the relationship between demolition and construction permit activities across Chicago block groups. If redevelopment is a primary source of housing supply, construction permit rates should be strongly correlated with demolition rates.

Figure 1 shows that demolition and construction activities are highly correlated across Chicago block groups. Regressing the demolition rate on the construction rate gives an OLS coefficient of 0.79 with an R^2 of 0.27 (Panel (a)). Roughly a third of block groups, most of which are low-income and in the South Side, have a positive demolition rate but no construction at all. In these locations, teardowns clear old, dilapidated buildings for public safety reasons. Panel (b) reports the same relationship using the redevelopment rate, which isolates true redevelopment events: the OLS coefficient is 0.80 with an R^2 of 0.87. The redevelopment rate almost perfectly correlates with the construction rate; across block groups, roughly four-fifths of the increase in construction is associated with the increase in redevelopment.

[8] The spatial concentration of housing redevelopment has also been documented in Munneke and Womack (2015) using data from the city of Miami.

[9] Among new-construction permits issued in 2009–2019 and matched to a parcel, 74% are on parcels with one to six assessed units. The unit share is from ACS 5-year table B25024: one-to-four-unit buildings and five-to-nine-unit buildings hold 59% and 11.8% of housing units in 2015–2019, respectively. The ACS does not subdivide the five-to-nine-unit category.



(a) Demolition and construction

(b) Redevelopment and construction

Figure 1: Demolition, Redevelopment, and Construction across Neighborhoods in Chicago, 2009–2019

Notes: The figure shows binned scatterplots of block-group permit rates in Chicago over 2009–2019. Panel (a) plots the demolition rate against the construction rate; Panel (b) plots the redevelopment rate, the share of parcels with a demolition permit matched to a new-construction permit on the same parcel, against the construction rate. Block groups are sorted by the construction rate into 40 equal-frequency bins; each dot is a bin mean, with dot size proportional to the number of parcels in the bin. Coefficients, standard errors, and R^2 are from OLS using block-group level data. The regressions are weighted by the number of parcels.

At the parcel level, among parcels of one to six residential units that receive a new-construction permit between 2009 and 2019, 63% are also associated with a demolition permit issued on the same parcel. The match rate is itself higher where construction is more active: it rises from 45% in the quartile of block groups with the lowest construction rates to 76% in the highest quartile, indicating that redevelopment plays a larger role in neighborhoods with more construction. Collectively, the block-group-level and parcel-level evidence points to redevelopment as a primary source of housing supply in Chicago. In a dense urban environment, one largely has to tear down an old building to build a new one.

Renovation rates are also related to demolition and construction rates,[10] though the regression coefficients are much smaller—0.25 and 0.28 (Figure F.2), compared with 0.80 between the redevelopment and construction rates—suggesting that renovation is an alternative way of upgrading the housing stock. We include renovation as an option available to landlords in the model.

[10]We construct the renovation rate from addition and remodel permits only, excluding permits involving minor repairs and deconversion.

Table 1: Permit rates on block group characteristics

Dependent variable	Permit rate			
	Demolition (1)	Construction (2)	Redevelopment (3)	Renovation (4)
Log household income	-0.32^{***}	-0.05^{***}	-0.04^{***}	-0.29^{***}
	(0.03)	(0.02)	(0.01)	(0.03)
Log gross rent	0.37^{***}	0.16^{***}	0.16^{***}	0.20^{***}
	(0.04)	(0.02)	(0.02)	(0.04)
Log home value	0.18^{***}	0.31^{***}	0.26^{***}	0.53^{***}
	(0.03)	(0.01)	(0.01)	(0.02)
Building age (decades)	-0.41^{***}	-0.25^{***}	-0.20^{***}	-0.49^{***}
	(0.07)	(0.04)	(0.03)	(0.06)
Building age squared	0.04^{***}	0.02^{***}	0.02^{***}	0.04^{***}
	(0.01)	(0.00)	(0.00)	(0.00)
Mean of dep. var.	0.22	0.11	0.08	0.41
Observations	2,331	2,331	2,331	2,331
R^2	0.09	0.32	0.33	0.30

Notes: Permit rates are the share of parcels receiving a permit of each type, averaged over 2009–2019; the redevelopment rate in column (3) counts demolition permits matched to a new-construction permit on the same parcel. Right-hand-side variables are block-group medians from the ACS five-year estimates over the same years. Regressions and the mean of the dependent variable are weighted by the number of parcels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

3.2 Redevelopment as a Quality-Upgrading Margin

We next examine how neighborhood characteristics shape redevelopment along both the extensive margin—the overall intensity of permit activity—and the intensive margin—the structural attributes of newly developed buildings.

Permit rates and neighborhood characteristics. On the extensive margin, we compare permit rates across block groups with different characteristics in Table 1. The table reports regressions of the four permit rates—demolition, construction, redevelopment, and renovation—on block-group-level characteristics, including median household income, median gross rent, median home value, and a quadratic in median building age, measured in decades, from the ACS. All variables are decadal averages over 2009–2019, and regressions are weighted by the number of parcels in the block group.

All four permit rates display qualitatively similar patterns. They correlate positively with housing rents and home values—consistent with building activity responding to local housing demand—yet negatively with neighborhood income. Building activities thus concentrate in neighborhoods with high housing costs but low income, precisely those most susceptible to gentrification.[11] The relationship between permit rates and median building age is U-shaped: rates first decline and then rise as the housing stock ages, turning upward at roughly fifty years for demolition, so the oldest neighborhoods experience the highest building intensity. Figure F.3 further shows that demolition rates are higher in tracts where land accounts for a larger share of property value, using the land share data produced by Davis et al. (2021). This pattern is consistent with redevelopment being more profitable where structures have low residual value relative to their sites.[12]

Parcel-level changes at redevelopment. In what follows, we characterize how redevelopment changes the housing stock and whether such changes differ across neighborhoods. To do this, we group parcels that are associated with subdivision or assembly over time; parcels with unchanging boundaries are assigned to their own group. We define a redevelopment event as a group of parcels that are 1) associated with a demolition permit and subsequent construction permit and 2) record a change in build year in the assessments. We measure the change in housing units and floorspace as the difference in totals across the whole parcel group.[13] Figure 2 plots the distributions of parcel group-level changes in units, log total floorspace, and log floorspace per unit at redevelopment events, separately for block groups below and above sample median income.

Redevelopment reduces unit counts while expanding total floorspace and floorspace per unit. Both margins are stronger in higher-income neighborhoods: unit counts fall by 0.16 in below-median-income block groups and by 0.59 in above-median ones, and floorspace per unit rises by 0.78 and 0.97 log points respectively, while total floorspace rises by 0.66 and 0.61 log points. New structures are larger—an important indicator of quality upgrading—but

[11] Table F.1 shows that these relationships are non-linear in income and housing costs, with permit rates tending to be higher in both low-income, low-housing-cost and high-income, high-housing-cost neighborhoods. Also, the negative demolition–income relationship is partly driven by extremely poor neighborhoods that experienced widespread teardowns without offsetting new construction after the Great Recession. Excluding these block groups leaves the rent and home-value patterns intact but eliminates the income relationship for demolition permits.

[12] We report the land share separately rather than adding it to Table 1: it is available for only 56% of our block groups, and the ones it misses have roughly twice the average demolition rate.

[13] A redeveloped parcel need not retain its assessment identifier as parcels are occasionally consolidated or subdivided upon redevelopment. To group parcels associated with the same consolidation or subdivision project, we overlay the 2005 and 2024 Cook County parcel maps. We define connected groups of 2005 and 2024 parcels linked whenever their overlap covers at least half of the smaller parcel. To ensure overlapping parcel groups correspond to the same consolidation or assembly project, we retain only groups whose 2005 and 2024 total areas agree to within a 15% tolerance. Units and floorspace are summed within each parcel group, so consolidating two single-family lots into one duplex leaves the unit count unchanged, while splitting one lot into two single-family homes adds a unit. See Appendix A.1 for details.

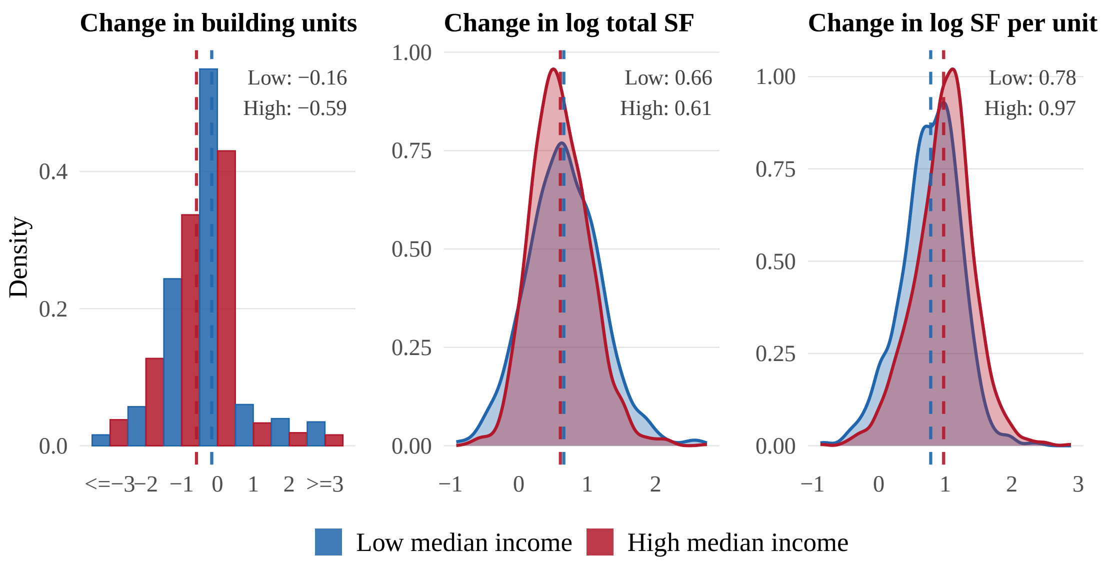


Figure 2: Distribution of changes in housing characteristics at redevelopment events.

Notes: Block groups are split into low versus high median income at the sample median. For change in building units, values outside $[-3, 3]$ are collapsed into the end bins; for changes in log total SF and log SF per unit, observations outside $[-1, 3]$ are excluded. Dashed lines indicate group means; means are shown in the top-right of each panel.

contain fewer units than the buildings they replace.[14] This quantity-quality tradeoff is more pronounced in high-income block groups, where unit counts fall further and floorspace gains are larger. Redevelopment thus operates primarily as a vehicle for quality upgrading at the cost of housing unit reductions, especially in higher-income neighborhoods.

One plausible source of this tradeoff is zoning regulation. Chicago's zoning code caps the floor-area ratio (FAR), the ratio of floorspace to lot area, on each parcel. Where the cap binds, landlords can only build large, high-quality housing by reducing the number of housing units. In neighborhoods where building larger units is more profitable than building more units, . In the model, we account for such land-use constraints by allowing construction costs to vary flexibly with quality and unit count and calibrate them to match revealed redevelopment choices. Building a high-quality multi-family structure can be prohibitively costly due to the zoning regulations.

In Table F.2, we regress the characteristics of new buildings, and their changes relative to the buildings they replace, on a broader set of block-group characteristics. The results confirm that the quantity-quality tradeoff is driven by demand-side conditions: unit counts fall further

[14]Average household size in Chicago has declined steadily over the past century: from 4.68 persons per household in 1900 to 3.47 in 1950, 2.67 in 2000, and 2.25 in 2020 (U.S. Decennial Census). Shrinking households, if anything, would increase the demand for smaller units, the opposite of what redevelopment delivers.

in higher-value neighborhoods, and floorspace per unit rises more in both higher-income and higher-value neighborhoods. Median building age, though correlated with the redevelopment rate, has no systematic effect on the characteristics of new buildings, suggesting that the vintage of the existing stock shapes when redevelopment occurs but not how. Consistent with our evidence, Institute for Housing Studies (2021) finds that, between 2013 and 2019, Chicago lost more than 4,800 two-to-four-unit buildings and that 47.5 percent of these lost buildings were replaced by single-family homes. These losses concentrate in high-rent neighborhoods and in moderate-rent areas with rising home values.

3.3 Redevelopment as a Gentrification Force

We now investigate the local impacts of redevelopment on neighborhood outcomes. Our main measure of the redevelopment rate is the minimum of the block-group demolition and construction permit rates over 2009–2019, expressed in percentage points. We prefer this measure because it captures the change in the building quality of the neighborhood through the replacement of old structures with new ones. We estimate the following regression:

$$\Delta Y_b = \beta_0 + \beta_1 \text{Redevelopment Rate}_b + \mathbf{X}_b' \boldsymbol{\delta} + \epsilon_b \tag{1}$$

where ΔY_b is the decadal log change (2009–2019) in outcome Y, including the number of housing units, median gross rent, median housing value, total population, and median income, of block group b, and $\mathbf{X}_b$ is a vector of controls. We restrict the sample to block groups in which more than 80% of housing units are initially in buildings with 1–4 units as of 2009. The restriction keeps the sample within the scope of the model and of the assessment records, both of which describe buildings of one to six units. Unobserved local demand shocks (e.g., amenity changes) and supply shocks (e.g., construction cost changes) may affect both redevelopment rates and local outcomes, which calls for an instrumental variable.

Our instrument is the interaction of a commuting-based shift-share labor demand shock with the initial share of pre-1910 housing at the block-group level. Our empirical design follows Diamond (2016), who applies the interaction between a Bartik-style labor demand shock and the housing supply elasticity as the instrument for changes in housing rent. Analogously, our identification exploits the fact that neighborhoods with larger shares of old housing should respond more strongly to positive demand shocks through redevelopment, since the low residual value of older buildings makes redevelopment more profitable. We construct the labor demand shock as the sum across destination block groups more than two kilometers away of the share of the block group's employed residents who worked there in 2002 times predicted employment growth at that destination between 2009 and 2019, where predicted growth is the destination's initial industry composition times national industry growth outside Illinois (Baum-Snow and Han, 2024). We assume that these national industry shifts are exogenous to neighborhood-level shocks in Chicago and exclude construction employment from the shift-share variable: its growth may reflect construction productivity shocks that shift housing supply. We use 1910 as the cutoff because the majority of demolished buildings in our sample were built before that year (Figure F.4).

We control for the commuting shock and for a vector of building vintage shares — those built before 1900, 1910, and so on through 1960 — so that the identifying variation comes from their interaction alone. The first absorbs the direct effect of local labor demand shocks on neighborhood outcomes; the second absorbs any trend common to older neighborhoods, as historic buildings and streets may carry amenity values. Following Baum-Snow and Han (2024), we also control for initial median income, which absorbs mean reversion in the local income process and land-use regulation that varies systematically with neighborhood income, and for a cubic polynomial in the block group's distance to the Loop, Chicago's central business district, which absorbs housing demand shocks that vary smoothly with centrality such as the growing attraction of central neighborhoods to high-income households. Conditional on these controls, the identification assumption is that the interaction affects neighborhood outcomes only through redevelopment.

First stage. The first-stage coefficient on the interaction between the commuting shock and the pre-1910 housing share is positive and significant (Table F.3): when demand rises, redevelopment responds most where the existing structures are oldest. The Kleibergen-Paap F-statistic is 25.80, above the conventional threshold of 10.

IV estimates. Table 2 reports the IV estimates.[15] Regressions are weighted by block-group population, and standard errors are clustered at the census tract level. Redevelopment raises neighborhood housing values and incomes: the coefficients on log housing value and log income are 0.10 and 0.16, both significant at the 1% level. A one-standard-deviation increase in the redevelopment rate raises the growth rates of housing value and income by 0.80 and 0.86 standard deviations. These estimated effects indicate that redevelopment is a strong force of neighborhood gentrification. However, neither the total number of housing units nor population increases significantly, consistent with the parcel-level evidence in Figure 2 that redevelopment replaces existing structures with fewer but larger units. The rent effect is small and imprecisely estimated: renters make up a small share of residents in these single-family-

[15]The number of observations is smaller for the rent regression because the ACS suppresses median gross rent for block groups with fewer than 10 sampled rental households, which is common in block groups with large shares of 1–4 unit housing.

Table 2: The effects of redevelopment on neighborhood outcomes

Dependent variable	Changes in				
	Units (1)	Rent (2)	Value (3)	Population (4)	Income (5)
Redevelopment rate	0.07	0.02	0.10***	0.07	0.16***
	(0.07)	(0.04)	(0.04)	(0.07)	(0.04)
Commuting shock	Yes	Yes	Yes	Yes	Yes
Initial log income	Yes	Yes	Yes	Yes	Yes
Vintage share controls	Yes	Yes	Yes	Yes	Yes
Cubic in distance to the Loop	Yes	Yes	Yes	Yes	Yes
Observations	1,277	1,039	1,277	1,277	1,277
Kleibergen-Paap F-stat	25.80	22.59	25.80	25.80	25.80

Notes: The redevelopment rate is the minimum of the block-group demolition and construction permit rates, averaged over 2009–2019 and expressed in percentage points, instrumented with the interaction of the commuting shift-share labor demand shock described in the text and the initial (2009) share of housing built before 1910. Dependent variables are 2009–2019 log changes. The sample is block groups in which more than 80% of housing units are initially, in 2009, in buildings with 1–4 units. Vintage share controls are the initial shares of housing built before 1900, 1910, 1920, 1930, 1940, 1950, and 1960; the distance control is a cubic polynomial in the distance from the block-group centroid to the Loop. Regressions are weighted by block-group population; standard errors clustered at the Census tract level in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

dominated neighborhoods, which leaves median gross rent noisily measured. Redevelopment is associated with improvements in local amenities: Table F.4 shows that block groups with a higher redevelopment rate also experienced decreases in crime and increases in the number of services establishments, though the latter estimate is not statistically significant. Therefore, redevelopment upgrades the quality of the existing stock, improves local amenities, and attracts higher-income households, all without expanding the supply of housing.

Robustness. Table F.4 reports the robustness checks on the housing-value and income estimates. The estimates are robust to constructing the instrument from another vintage cutoff year in the early 1900s, e.g. 1920, and to relaxing the sample restriction to block groups with an initial 1–4 unit share above 60%. Including block groups dominated by large buildings yields somewhat smaller but still significant estimates. Finally, the IV estimates are robust to measuring the redevelopment rate by the demolition rate alone or by the rate of linked demolition and construction permits (Panels C and D of Table F.4). The estimated effects of the demolition rate are smaller because in many areas demolition is not followed by construction.

3.4 Summary of the Empirical Facts

We have established three empirical facts about redevelopment in this section. First, redevelopment is a primary source of housing supply in dense urban neighborhoods and concentrates where incomes are low, rents are high, and the housing stock is old. Second, redevelopment is a quality-upgrading margin: it shifts the stock toward fewer but larger units. Third, redevelopment is a force of gentrification: it causally raises neighborhood housing values and incomes without appearing to expand net unit supply or population. Together, these facts call for a model in which redevelopment and income sorting are determined jointly. In the next section, we build such a model.

4 Model

We now build a dynamic urban model of housing redevelopment and income sorting. Consider a city composed of neighborhoods $x \in \mathbb{X}$ and an outside option o with utility normalized to one. Neighborhoods differ in income-specific amenities $\bar{A}(x, z)$, and the neighborhood amenities also respond to their average income. Each neighborhood contains parcels $i \in \mathbb{I}_x$ owned by immobile landlords. Each parcel contains a residential building with housing quality $q \in \mathbb{Q}$ and a number of units $h \in \mathcal{H} = \{0, 1, 2, \ldots, H\}$, where $h = 0$ denotes an empty parcel. Quality is a one-dimensional index that summarizes a building's characteristics, including its age, floor space, construction materials, and heating and cooling systems and so on. Households ω differ in income $z_\omega \in [0, Z_{\max}]$ and origin $x_{\omega_0} \in \mathbb{X} \cup \{o\}$, with exogenous measure $\bar{L}(z, x_0)$. We assume in the model that every households rent a housing unit and treat an owner-occupier as a household that rents from itself.[16] Time is discrete and indexed by t. Housing quality depreciates at rate δ each period, and both landlords and households are forward-looking. We omit the time index t and household index ω when no confusion arises.

4.1 The Household's Problem

The household's problem consists of two choices: which neighborhood to live in, and what quality of housing to rent within it. We present the problem in reverse order.

Housing quality choice

A household with income z_ω in neighborhood x chooses housing quality q and the numeraire y to maximize

$$U(x, z_\omega) = \max_{q,y} (q - \bar{q})^\alpha y^{1-\alpha} \tag{2}$$

[16] Absent taxes and transaction costs, the user cost of owning a unit equals its market rent Poterba (1984).

subject to

$$P(q,x) + y = z_\omega \tag{3}$$

where $P(q,x)$ is the equilibrium rent for quality q in neighborhood x, α governs the housing expenditure share, and $\bar{q} > 0$ is the minimum quality demanded by households. The parameter $\bar{q}$ governs how the housing expenditure share declines with income and thus the degree of non-homotheticity in housing demand (Albouy et al., 2016; Couture et al., 2024). Assuming that the rent function is differentiable, the optimal quality choice satisfies

$$\underbrace{\frac{\alpha}{1-\alpha}\frac{z_\omega - P(q,x)}{q-\bar{q}}}_{\text{MRS of housing for numeraire}} = \underbrace{\frac{\partial P(q,x)}{\partial q}}_{\text{Slope of rent function}} . \tag{4}$$

Equation (4) states that households choose quality until their willingness to pay for a marginal quality increase equals the associated rent increase. For any strictly increasing rent function $P(q,x)$, the MRS is strictly decreasing in q, so higher-income residents match to strictly higher-quality housing within each neighborhood (Määttänen and Terviö, 2014). The rent function $P(q,x)$ is the central equilibrium object: it shapes income sorting across neighborhoods and, through landlords' decisions, the quality of new housing supply. In Section 4.4, we formalize these two roles.

Neighborhood choice

$$V_{\omega t}(z_\omega, x_{\omega_0}, \vec{\epsilon}_t) = \max_x \ \tau(x; x_{\omega_0}) \cdot U_t(x, z_\omega) A_t(x, z_\omega)\epsilon_{\omega t}(x) + \beta \mathbb{E}_{\vec{\epsilon}}\, V_{t+1}(z_\omega, x_{\omega_0}, \vec{\epsilon}_{\omega t+1}), \tag{5}$$

where x_{ω_0} is the household's original neighborhood, $\tau(x; x_{\omega_0})$ is the mobility cost of moving to x, $A(x,z)$ is the income-specific amenity of neighborhood x, $\vec{\epsilon}_{\omega t} = (\epsilon_{\omega t}(1), ..., \epsilon_{\omega t}(X), \epsilon_{\omega t}(o))$ is a vector of neighborhood-specific idiosyncratic preference shocks, and β is the discount factor.

We make three assumptions. First, following Guerrieri et al. (2013), neighborhood amenities respond endogenously to average income:

$$A_t(x,z) = \bar{A}(x,z) \cdot \bar{z}_t(x)^\eta, \tag{6}$$

where $\bar{A}(x,z)$ is the exogenous, income-specific amenity component, $\bar{z}_t(x)$ is average neighborhood income, and η is the amenity spillover elasticity. Second, moving costs take the form of a utility shifter: $\tau(x; x_{\omega_0}) < 1$ if $x \neq x_{\omega_0}$ and $\tau(x;x) = 1$ for all x, capturing preferential attachment to the origin neighborhood. As in Desmet, Nagy and Rossi-Hansberg (2018), this formulation renders the neighborhood choice problem static, since households can always return to their origin in a future period. Third, the idiosyncratic preferences $\{\epsilon_{\omega t}(x)\}_{\forall \omega, t, x}$ are drawn i.i.d. from a Type-II Extreme Value distribution with dispersion parameter σ_x, i.e., $F(\epsilon) = \exp\left(-\epsilon^{-\sigma_x}\right)$.

These assumptions yield a closed-form expression for the mass of type-(z, x_0) households choosing neighborhood x:

$$L_t(x, z, x_0) = \bar{L}(z, x_0) \cdot \frac{\left[\tau(x; x_0) \cdot U_t(x, z) A_t(x, z)\right]^{\sigma_x}}{\sum_{x' \in \mathbb{X}} \left[\tau(x'; x_0) \cdot U_t(x', z) A_t(x'; z)\right]^{\sigma_x} + \tau(o; x_0)^{\sigma_x}}, \tag{7}$$

where the last term in the denominator captures the outside option with normalized utility of 1. Both margins of the household problem take the rent function as given. In equilibrium, $P(q, x)$ is determined jointly by household demand and the supply of housing quality. We now turn to the landlord's redevelopment problem, which governs the supply side.

4.2 The Landlord's Problem

Each landlord maximizes the discounted value of its parcel. Let $s_{it} = (q_{it}, h_{it})$ denote the state of parcel i in period t: housing quality q and unit count $h \in \mathcal{H} = \{0, 1, 2, \ldots, H\}$, where $h = 0$ corresponds to an undeveloped parcel. At the beginning of each period, the landlord chooses among three actions: (i) retain the existing structure as its quality depreciates (N); (ii) renovate, raising quality while preserving the unit count (RN); or (iii) redevelop by replacing the existing structure with a new one under the quality blueprint and choosing the new structure's number of units (R).

Let $V_{it}(s_{it}, \hat{q}_{it}, \vec{\xi})$ denote the value of parcel $i \in \mathbb{I}_x$, where $\hat{q}_{it}$ is a redevelopment quality blueprint (defined below) and $\vec{\xi}_{it}$ collects idiosyncratic option-specific profit shocks. The landlord's problem is formally defined as

$$V_{it}(s_{it}, \hat{q}_{it}, \vec{\xi}_{it}) = \max\left\{ \; V_{it}^{N}(s_{it}) + \frac{1}{\sigma_c}\xi_{it}^{N}, \; V_{it}^{RN}(s_{it}) + \frac{1}{\sigma_c}\xi_{it}^{RN}, \; \max_{h \in \mathcal{H}} \left[V_{it}^{R}(\hat{q}_{it}, h) + \frac{1}{\sigma_c}\xi_{it}(h) \right] \; \right\}, \tag{8}$$

where ξ_{it}^{N}, ξ_{it}^{RN}, and $\{\xi_{it}(h)\}_{h \in \mathcal{H}}$ are drawn i.i.d. from a Type-I Extreme Value distribution with its dispersion scaled by parameter σ_c, which we assume to be constant across neighborhoods. The parameter σ_c thus governs how the redevelopment and renovation decisions respond to changes in the associated returns. Ma (2025) studies developers' parcel-level choice of whether and what quality type of single-family home to build. We extend her supply model by allowing landlords to choose the number of units and incorporating renovation, while allowing for quality choice through intertemporal selection among blueprint opportunities. We next describe each option in turn.

Inaction and renovation. Under inaction (N), the landlord collects rent while housing quality depreciates at rate δ. Define the depreciation operator $D(q) = \max\{(1-\delta)q,\, \bar{q}\}$, where structures at the minimum quality $\bar{q}$ are effectively vacant. The value of inaction is

$$V_{it}^{N}(s_{it}) = P_t(q_{it}, x)h_{it} \, + \beta \mathbb{E}_{\hat{q}_{it+1},\, \vec{\xi}_{it+1}} V_{i,t+1}\left(D(q_{it}), h_{it}, \hat{q}_{i,t+1}, \vec{\xi}_{it+1}\right).$$

Under renovation (RN), the landlord pays the renovation cost $C_x^{RN}(q,h)$, to raise quality on all units:

$$V_{it}^{RN}(s_{it}) = -C_x^{RN}(q_{it}, h_{it}) + P_t\big((1+\varrho)q_{it}, x\big)h_{it} + \beta \mathbb{E}_{\hat{q}_{it+1},\, \vec{\xi}_{it+1}} V_{i,t+1}\left(D\big((1+\varrho)q_{it}\big), h_{it}, \hat{q}_{i,t+1}, \vec{\xi}_{it+1}\right),$$

where we assume that renovation improves all units' quality proportionally by ϱ. We abstract from the landlord's choice of how much quality to add through renovation.

Redevelopment. Each period, the landlord draws a redevelopment blueprint $\hat{q}_{it}$ from a distribution $G(\hat{q})$, where $\hat{q} \in \mathbb{Q}' \subseteq \mathbb{Q}$. The random blueprint captures, in reduced form, matching frictions between landlords and heterogeneous construction firms and the inability of the landlord to perfectly arbitrage between different quality segments (Damen et al., 2025; Greenwald and Guren, 2025). It gives the landlord the option—but not the obligation—to redevelop at quality $\hat{q}_{it}$ and choose the number of units h. If the landlord instead chooses inaction or renovation, the blueprint expires and a new one is drawn next period. The value of building h units at quality $\hat{q}_{it}$ is

$$V_{it}^{R}(\hat{q}_{it}, h) = P_t(\hat{q}_{it}, x)h - C_x(\hat{q}_{it}, h) + \beta \mathbb{E}_{\hat{q}_{i,t+1}, \vec{\xi}_{i,t+1}} V_{i,t+1}\left(D(\hat{q}_{it}), h, \hat{q}_{i,t+1}, \vec{\xi}_{i,t+1}\right).$$

Because redevelopment replaces the existing structure, which has no salvage value, this value does not depend on the current state s_{it}. We assume that the construction sector is perfectly competitive, a common assumption in the housing production literature (Epple et al., 2010; Combes et al., 2021), so the landlord captures the redevelopment surplus. Redevelopment costs take the form

$$C_x(\hat{q}, h) = \underbrace{\Omega_x(h)\hat{q}}_{\text{Variable cost}} + \underbrace{F_{\hat{q}x}}_{\text{Fixed cost}}, \tag{9}$$

where the variable cost shifter $\Omega_x(h)$ maps the number of units into variable cost per quality unit in neighborhood x, embedding construction technology and the shadow cost of zoning regulations; the fixed cost $F_{\hat{q}x}$ varies across blueprint qualities and neighborhoods and covers the physical costs of teardown and site preparation as well as regulatory costs, including the teardown tax.

Choice probabilities. Given the properties of the Type-I Extreme Value distribution, we can solve for the choice probabilities of renovation and redevelopment as

$$P_{it}^{RN}(s_{it},\hat{q}_{it}) \;=\; \frac{\exp\left(\sigma_c V_{it}^{RN}(s_{it})\right)}{\tilde{V}_{it}(s_{it},\hat{q}_{it})} \quad ; \quad P_{it}^{R}(s_{it},\hat{q}_{it}) \;=\; \frac{\sum_{h'\in\mathcal{H}}\exp\left(\sigma_c V_{it}^{R}(\hat{q}_{it},h')\right)}{\tilde{V}_{it}(s_{it},\hat{q}_{it})}, \tag{10}$$

respectively, where $\tilde{V}_{it}(s_{it},\hat{q}_{it}) \equiv \exp\left(\sigma_c V_{it}^{N}(s_{it})\right)+\exp\left(\sigma_c V_{it}^{RN}(s_{it})\right)+\sum_{h'\in\mathcal{H}}\exp\left(\sigma_c V_{it}^{R}(\hat{q}_{it},h')\right)$. Conditional on redevelopment, the probability of building h units is

$$P_{it}(h \mid \hat{q}_{it}) \;=\; \frac{\exp\left(\sigma_c V_{it}^{R}(\hat{q}_{it},h)\right)}{\sum_{h'\in\mathcal{H}}\exp\left(\sigma_c V_{it}^{R}(\hat{q}_{it},h')\right)}. \tag{11}$$

4.3 Equilibrium

We assume that landlords have perfect foresight over future rent schedules across the entire quality distribution.[17]

Definition 1 *Given initial housing conditions $\{s_{i0}\}_{\forall i}$ and the income distribution by origin $\{\bar{L}(z,x_0)\}_{\forall z,x_0}$, a perfect foresight equilibrium is a set of housing prices $\{P_t(q,x)\}_{\forall t,q,x}$, housing quality distributions $\{H_t(q,x)\}_{\forall t,q,x}$, quality and neighborhood choices $\{L_t(q,x,x_0,z)\}_{\forall q,x,x_0,z,t}$, and landlord value functions $\{V_{it}(s,\hat{q},\vec{\xi})\}_{\forall i,t,s,\hat{q},\vec{\xi}}$ such that in each period t:*

1. *Each household chooses housing quality and neighborhood to maximize utility according to (2) and (5).*

2. *Each landlord makes renovation and redevelopment decisions to maximize the value of its parcel, according to (8).*

3. *Housing markets clear in all neighborhoods and quality types:*

$$\int_{z\in\mathbb{Z}_t(q,x)} L_t(q,x,z)\,dz = H_t(q,x), \quad \forall q>\bar{q},x,t, \tag{12}$$

   where $\mathbb{Z}_t(q,x)$ is the set of incomes choosing q in x at t, $H_t(q,x)\equiv\sum_{i\in\mathbb{I}_x}\mathbb{1}\{q_{it}=q\}\cdot h_{it}$, and $L_t(q,x,z)\equiv\sum_{x_0\in\mathbb{X}\cup\{o\}}L_t(q,x,x_0,z)$.

4. *The rent schedules that landlords foresee coincide with the equilibrium paths.*

[17] We abstract from belief formation and coordination in housing redevelopment, which are studied by Murphy (2018) and Owens et al. (2020).

We solve for the equilibrium sequence of rents and allocations as a fixed point in the path of rent functions. Given a candidate path that landlords take as given, we solve their problems by backward induction, simulate the implied renovation and redevelopment decisions forward to construct the housing stock in each period, and update rents to clear the market for each neighborhood, quality segment, and time. We iterate until the rent path converges. Appendix B.2 provides implementation details.

We apply this model to understand the dynamic effects of various housing policies that impact both the quality and quantity of new housing. To understand the model's theoretical predictions about these effects, we analyze a simplified, static version of the model in the following section.

4.4 Comparative Statics: Income Sorting, Redevelopment, and Anti-Redevelopment Policies

The rent function is an important equilibrium object: it determines the quality and quantity of the housing stock by shaping landlords' incentives to redevelop as well as neighborhood income in spatial equilibrium. In this subsection, we provide analytical results that show the rent function's impact on redevelopment and income sorting in partial equilibrium. In turn, we assess how exogenous changes in a neighborhood's housing quality and quantity pass through to prices, determining the equilibrium structure of the rent function. The formal statements and proofs of the theoretical results in this subsection can be found in Appendix B.

Rent elasticity, sorting, and redevelopment. We first analyze how the rent elasticity affects income sorting and redevelopment. In order to establish simple theoretical results, we work in a one-period, partial equilibrium setting in which households have Cobb–Douglas preferences ($\bar{q} = 0$) and zero moving costs, landlords choose only between inaction and redevelopment at a single blueprint quality, and rents are iso-elastic in quality, $P(q, x) = \kappa(x)q^{\nu(x)}$; Assumption 1 in Appendix B states the environment formally.

Proposition 1 (Rent elasticity, sorting, and redevelopment) *Under Assumption 1, compare two neighborhoods whose rent functions have elasticities $\nu_2 > \nu_1 > 0$, so that high-quality housing is relatively more expensive in neighborhood 2.*

1. Income sorting. *The indirect utility of neighborhood 1 relative to neighborhood 2 is strictly increasing in household income. Higher-income households therefore sort into neighborhood 1, which has the higher average income.*

2. Redevelopment. *Holding quality and the number of units constant, a parcel in neighborhood 2 has the higher redevelopment rate for sufficiently high quality blueprints.*

Part 1 reflects how the elasticity of the rent schedule prices income. When the schedule is steep, each additional dollar of housing spending buys less quality, so the utility gain from additional income is smaller; high-income households therefore value the neighborhood relatively less. As a result, when housing is indivisible and each household rents one unit, variation in the rent function can generate income sorting across space even when preferences are homothetic—a channel absent from standard models with divisible housing (Diamond and Gaubert, 2022). Moreover, the addition of non-homothetic preferences, i.e., $\bar{q} > 0$, makes the level of the rent schedule matter for income sorting as well; this fact is well-studied in the literature (Couture et al., 2024; Finlay and Williams, 2025). Part 2 turns to the supply side. The landlord's incentive to redevelop is governed by the quality premium in rent. Where high quality is expensive relative to low, rebuilding at high quality commands a larger rent gain, so of two parcels with the same quality and unit count, the one in the steeper neighborhood is redeveloped more often once the blueprint quality is high enough.

Proposition 1 sheds light on the joint dynamics of redevelopment and income sorting. Holding neighborhood amenities constant, where high-quality units are scarce, the rent function is steep, so average income is low (Part 1) and redevelopment is likely (Part 2).[18] Redevelopment adds high-quality units, flattens the schedule, and draws in high-income households, as Table 2 showed.[19] We test the proposition's predictions in Section 5, correlating the estimated rent elasticity with the redevelopment rate and the change in neighborhood income.

Housing quality, quantity, and rental affordability. We now ask how changes in a neighborhood's housing quality and quantity pass through to rents and income sorting.

Proposition 2 (Quality, quantity, and rental affordability) *Consider a neighborhood that is small relative to the city, and suppose its housing stock changes as follows, where the changes and the migration elasticity $\sigma_x > 0$ are sufficiently small.*[20]

[18] In the full model, amenities also shape the rent function: neighborhoods whose amenities attract high-income households face stronger demand for high quality, which steepens the schedule. Therefore, the rent elasticity need not correlate negatively with neighborhood income in the data. We examine their empirical relationship in Section 5.1.

[19] Rosenthal (2008) documents the empirical pattern of gradual housing quality decline and housing renewal, with the housing stock driving income composition. Our model builds on his insight. We endogenize the rent schedule over quality as an equilibrium object, which affects both redevelopment and income sorting.

[20] In the proposition, distributional shifts are in the sense of first-order stochastic dominance, and rent changes are relative to an ex-ante identical neighborhood whose housing stock is unchanged. Formally, under Assumption 2 in Appendix B, which collects the regularity conditions, there exists $\bar{\sigma} > 0$ such that both parts hold whenever $0 < \sigma_x < \bar{\sigma}$.

1. Quality shift. *If the quality distribution shifts down, with the number of units unchanged, rents fall below a threshold quality and rise above it. The neighborhood also becomes poorer.*

2. Quantity expansion. *If the number of housing units increases, with the quality distribution unchanged, rents fall at every quality.*

Two forces are at work in part 1. The first is the trickle-down force of the assignment model (Nathanson, 2025): holding residents fixed, the quality shift rematches each unit to a higher-income household, who bids up its rent, and this force accumulates up the ladder as high-quality units grow scarce. The second is household mobility, which also reshapes who lives in the neighborhood: the quality shift steepens the rent schedule, so by Proposition 1 high-income households find the neighborhood less attractive and move out; the weaker demand lowers the rent level, and the cheaper housing draws in low-income households. The mobility force dominates at the bottom of the ladder and the trickle-down force at the top, and in the full model endogenous amenities amplify the rent decline as high-income residents leave. Part 2 is the standard result that a housing supply expansion lowers rents.[21] When quality deteriorates but quantity increases—the intended outcome of the anti-redevelopment policies in Chicago that we discuss next—the added units push rents down throughout the quality distribution while the quality shift pushes rents up only at the top, and a large enough quantity expansion leaves the neighborhood cheaper at every quality.

Figure 3 illustrates both parts of Proposition 2 and further incorporates endogenous amenities. The top row simulates a downward shift in housing quality that corresponds to part 1. The no-mibility scenario isolates the direct effect: rents rise at every quality as all households downgrade in quality and bid up rents along the distribution. With migration, high-income households leave and the neighborhood grows poorer, so low-quality rents fall below baseline while scarce high-quality units remain expensive. Endogenous amenities further amplify income sorting and shift the rent function further down. The bottom row simulates an increase in the housing stock that corresponds to part 2. In this case, we cannot restrict migration because some units will be unoccupied. Rents fall across the entire quality ladder and households, disproportionately the low-income ones who are more sensitive to rent changes, migrate into the neighborhood. As a result, a quantity expansion in a neighborhood lowers its average income, and through endogenous amenities the shift in income distribution further shifts down the rent function.

[21] The size of the rent decline depends on the migration elasticity σ_x: the higher the elasticity, the more of the new supply is absorbed by in-migration from the rest of the city and the less the rent decline (Anenberg and Kung, 2020).

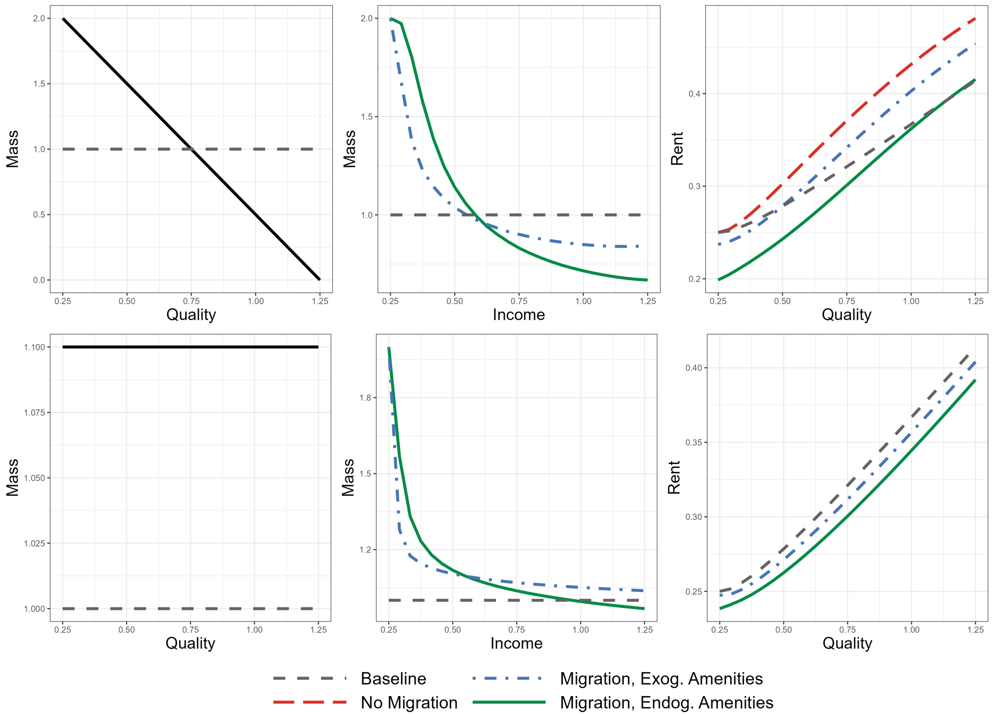


Figure 3: Illustrating the Equilibrium Effects of Housing Quality and Quantity Changes

Note: An assignment model with one neighborhood and an outside option neighborhood that provides exogenous utility. The baseline equilibrium (gray dashed) has uniform quality and income on the interval $[0.25, 1.25]$. In both rows, we simulate a change in the housing stock, represented by the solid lines. The no-mobility economy (red dashed) fixes residents at baseline; the mobility economy (blue dash-dotted) allows moves with amenities fixed; the full equilibrium (green solid) adds endogenous amenity responses to migration. Panels report the resident income distribution and the rent function across the four scenarios.

The two anti-redevelopment policies through the lens of the model. We use these theoretical results to analyze two anti-redevelopment policies implemented in two neighborhoods in Chicago: a teardown tax of T for the issuance of demolition permits and an anti-deconversion ordinance banning redevelopment that reduces the number of units. Both policies shift the neighborhood's quality distribution down by slowing redevelopment, while the anti-deconversion ordinance also aims to prevent the loss in housing units. By Proposition 2, the teardown tax lowers the rent of low-quality housing and curbs gentrification by making the high-quality housing less affordable, and the anti-deconversion ordinance's effect on preserving units would reduce rent across the quality distribution.

In this section, we used a simplified model to make sharp qualitative predictions about the effects of anti-redevelopment policies. In the following section, we calibrate the full dynamic model to make quantitative predictions about the effects of these policies and to do welfare analysis.

5 Estimation of Model Parameters

We estimate the model by matching salient features of housing quality and rent distributions, income compositions, and redevelopment activities across neighborhoods in Chicago. In particular, we exploit the anti-redevelopment policies implemented in two Chicago neighborhoods to identify the redevelopment elasticity. Our unit of analysis is a neighborhood group: we spatially cluster the 98 Chicago neighborhoods[22] into 25 groups with roughly 50,000 housing units each, and refer to these groups as "neighborhoods" hereafter (Table F.5). Two of these groups have housing stocks dominated by high-rise apartment buildings, which fall outside the scope of the model, so we estimate and solve the model on the remaining 23 groups. We describe the estimation procedure step by step below.

5.1 Estimating Rent Functions and the Depreciation Rate

We first use the hedonic regression to estimate a housing unit-year panel of housing quality, the depreciation rate, the renovation premium, and neighborhood-specific rent functions. Assume that housing quality is a uni-dimensional index that is log-additive in both observed and unobserved components:

$$\log q_{it} = \underbrace{-\,\delta \times \text{Building Age}_{it}}_{\text{Depreciation}} + \underbrace{\sum_{c\in\mathcal{C}} \alpha_c X^c_{it}}_{\text{Other observed characteristics}} + \underbrace{\log \epsilon^q_{it}}_{\text{Unobserved characteristics}}, \tag{13}$$

where i indexes housing unit, t indexes year, δ is the depreciation rate, $\mathcal{C}$ is the set of observed housing characteristics,[23] $X^c, c \in \mathcal{C}$ is the value of each characteristic, and ϵ^q is unobserved quality. We estimate neighborhood-specific rent functions, which we assume to be iso-elastic:

$$\log P(q, x) = \log \kappa(x) + \nu(x) \log q, \tag{14}$$

where κ governs neighborhood-specific rent level, and ν is the neighborhood-specific rent elasticity of housing quality. Substituting (13) into (14) yields a hedonic regression:

$$\log P_{it} = \log \kappa(x) + \nu(x) \left[-\delta \times \text{Building Age}_{it} + \sum_{c\in\mathcal{C}} \alpha_c X^c_{it}\right] + \nu(x) \log \epsilon^q_{it}, \tag{15}$$

[22] The neighborhood boundaries are from the Chicago Data Portal, available here.

[23] The characteristics include unit floor space, bedrooms and bathrooms per unit (with squared terms), land parcel size, number of floors, building type (single-family, small multi-family with four or fewer units, and medium multi-family), dummies for types of construction material, heating system, porch and garage, and an indicator for recent renovation permits.

where P represents housing rent. We estimate equation (15) using the method of nonlinear least squares, normalizing the average quality elasticity to one: $\frac{1}{|X|}\sum_{x\in X}\nu(x) = 1$.[24] Given the rent function parameter estimates, we can recover estimated housing quality as $\log \hat{q}_{it} = (\log P_{it} - \log \kappa(x))/\nu(x)$.

When estimating equation (15), we demean all variables by year to absorb city-level annual shocks in the housing market. We also exclude two neighborhoods—Millennium Park and River North–Gold Coast—whose housing stocks are dominated by high-rise apartment buildings outside the scope of the model, and buildings built before 1930, whose survival past the typical service life likely reflects exceptional maintenance or architectural premia, either of which would bias the depreciation estimate downward.[25]

Imputing market rents. Listed rental units are a selected subset of the housing stock, while the model requires a quality index for every unit. We therefore impute a market rent for every assessed property, adapting the rank-based procedure of Diamond and Diamond (2024). On the sample of listings linked to assessment records by address, we estimate how a unit's within-year rent rank varies with its within-year assessed-value rank and the characteristics in $\mathcal{C}$, apply the estimated mapping to all assessed properties, and convert predicted ranks into dollars using that year's distribution of listed rents. The identifying assumption is that, conditional on $\mathcal{C}$, the mapping from assessed-value rank to rent rank is the same for listed and unlisted units. Appendix A.2 details the imputation procedure.

Estimation result. Table F.6 reports the full set of hedonic coefficients. The signs align with expectations—larger units, higher construction quality, and single-family structures raise housing quality—and the observed characteristics fit the rent variation closely, with an R^2 of 0.94. The quality depreciation rate is $\delta = 0.34\%$ per year, close to the 0.4% estimate of Rosenthal (2014). The renovation indicator is estimated at 0.23 log points, so a major renovation raises housing quality by 26 percent ($\varrho = \exp(0.23) - 1$), roughly one standard deviation of estimated quality.

Our estimation reveals significant variation in the two rent function parameters across neighborhoods (Figure F.5). High rents arise for different reasons: the Near North Side pairs a low $\kappa(x)$ with a steep $\nu(x)$, the northern neighborhoods the reverse. The gradient rises with average income and with housing-stock age, while the level does not significantly vary by these local characteristics (Table F.7). Proposition 1 predicts that neighborhoods

[24]Because the quality index q has no cardinal scale, the normalization anchors quality so that a neighborhood with $\nu = 1$ has rent proportional to q.

[25]The BEA specifies a service life of 80 years for 1–4 unit residential structures. We recover quality for these pre-1930 units using the estimated parameters of (15).

with flatter rental functions should have high average income, holding amenities fixed. In the data, however, amenities are not held fixed. High-income households sort into neighborhoods with the amenities they value, and their demand for quality steepens the rent schedule there. Steeper rent functions are nonetheless associated with higher demolition rates and faster income growth between 2010 and 2019 (Figure F.6), consistent with Part 2 of Proposition 1.

5.2 Calibrating Housing Demand and Supply Parameters

We calibrate the remaining parameters in three steps: the housing preference parameters are chosen to match the distribution of rent expenditure across income groups, the redevelopment elasticity σ_c to match the estimated effects of the 606–Pilsen anti-gentrification policies on demolition, and the cost parameters that govern redevelopment and renovation to match the respective permit rates and improvement on housing quality. We also take a few parameters directly from the literature and discuss the calibration results in Section 5.3.

5.2.1 Period Length and the Quality Grid

For computational tractability, we set one model period to ten years and discretize the quality space coarsely.[26] We set a grid with ten quality levels; each gridpoint being equidistant from others in percentage terms. The grid spans quality five standard deviations from the average quality estimated from the hedonic regression (15). The lowest level is the minimum quality $\bar{q}$, at which structures provide no value and are effectively vacant.[27]

5.2.2 Demand-Side Parameters

We choose the housing preference parameters $(\alpha, \bar{q})$ to target rent expenditure shares across 7 household income groups—calculated using the American Community Survey (ACS)—and those implied by our model under the empirically estimated rent function $P(q, x)$.[28] The expenditure share function depends on both parameters: $\bar{q}$ governs by how much rent expenditure shares decrease with income, while α determines the average expenditure shares.

[26] The 10-year model period is unproblematic because virtually no parcels demolish more than once in this window of time. Moreover, our estimated depreciation rate is slow, so the dynamic effects of the policies we evaluate span multiple decades.

[27] The procedure yields a grid with adjacent quality levels that are 22 log points apart. However, because quality falls by only 3.3 percent per decade at the estimated annual rate of 0.34 percent, we model depreciation as stochastic. This means that each period a structure falls one grid level in quality with probability 0.15, and otherwise keeps its quality, so that the expected decline in log quality matches the estimated rate. We acknowledge that, due to convexity or concavity in the value function for land, stochastic depreciation may not lead to exactly the same redevelopment choices as with deterministic depreciation and the gap should narrow with finer quality grids.

[28] These income groups are defined as, in household yearly income: 0-25k, 25-50k, 50-75k, 75-100k, 100k-150k, 150-200k, 200k+. Rent expenditure shares fall from 49 percent for households earning below $25k to 12 percent for those above $200k.

We then invert the housing quality distribution $H(q,x)$ of each neighborhood from the solution to the household's quality choice problem under the estimated rent function and ACS neighborhood income distributions.[29]

5.2.3 Supply-Side Parameters

The supply-side parameters include the redevelopment elasticity σ_c, the fixed cost schedule $\{F_{\hat{q},x}\}$, the variable cost schedule $\{\Omega_x(h)\}$, the renovation cost C_x^{RN}, and the number of parcels by q and h. We now describe how these parameters are calibrated in turn.

Redevelopment Elasticity

We identify σ_c using quasi-experimental variation generated by a set of anti-gentrification policies designed to curb redevelopment in Chicago. In the decade before the policy, the 606 Trail and Pilsen neighborhoods (Figure D.1), which together account for roughly 4 percent of Chicago's housing stock, experienced much greater demolition activity than citywide averages, faster growth in rents and property prices (Figure D.2), and the conversion of multifamily buildings into single-family homes. In response, the City Council adopted two pilot ordinances for these areas, in force from April 2021 to April 2024: the Demolition Permit Surcharge, the greater of \$15,000 per permit and \$5,000 per demolished residential unit; and the Anti-Deconversion Ordinance, barring any reduction in a parcel's unit count.[30] Appendix D provides more institutional details.

We estimate the effect of the policy bundle on parcel-level redevelopment activity with two difference-in-differences designs. The first compares parcels within 500 meters inside the policy boundary to parcels within 500 meters outside with the following regression in the logistic form:

$$\log\left(\frac{\Pr(Y_{it}=1 \mid X_{it})}{1-\Pr(Y_{it}=1 \mid X_{it})}\right) = \delta_0 + \sum_{k=-t_0}^{t_1} \delta_k \, \mathbf{1}\{t=k\} \times Trt_i + \mu_i + \alpha_{xt} + F_t(LON_i, LAT_i), \quad (16)$$

where Y_{it} indicates that parcel i receives a demolition permit in three-year period t, $Trt_i = 1$ if parcel i is in the treated area and 0 otherwise, μ_i is the parcel fixed effect,[31] α_{xt} is an

[29] We target $H(q,x)$ generated by household quality choices in the model rather than aggregating the unit-level quality estimates of Section 5.1. We do this so that the baseline housing stock clears the market at the estimated rents—making the baseline an exact equilibrium of the model—and so that measurement error that affects quality estimation does not propagate into the cost parameters. Section 5.4 shows that across-neighborhood variation in quality that the model generates is close to the data.

[30] The surcharge was subsequently extended and later renewed with broader neighborhood coverage and a surcharge four times the original amount. We study a policy expansion in the counterfactual analysis of Section 6.

[31] A logit regression with parcel fixed effects is identified only from parcels on which a permit is issued over the sample period. Dropping the fixed effects and retaining every parcel leaves the demolition and construction estimates close to those above (Table D.4).

area-by-period fixed effect, where x indexes either 606 Trail or Pilsen and their respective buffer zones outside the treatment boundary, and $F_t(LON_i, LAT_i)$ interacts period dummies with longitude and latitude to control for time-varying spatial demand shocks across the city. The geographical controls ensure that the identification comes from the changes in demolition activities across the border. Parcels are defined according to the 2005 assessment files (see Section 2 for details).

The second design matches each treated block group with up to three untreated block groups at least one kilometer from the treatment areas and estimates (16) on the matched sample, dropping α_{xt} because these control block groups lie outside the areas x indexes.[32] The designs are complementary: the boundary design compares parcels with similar housing characteristics and demand conditions, but its control ring is more exposed to the policy's spillovers, though these appear limited,[33] while the matched control neighborhoods may face different local shocks even conditional on the time-varying geographic controls.

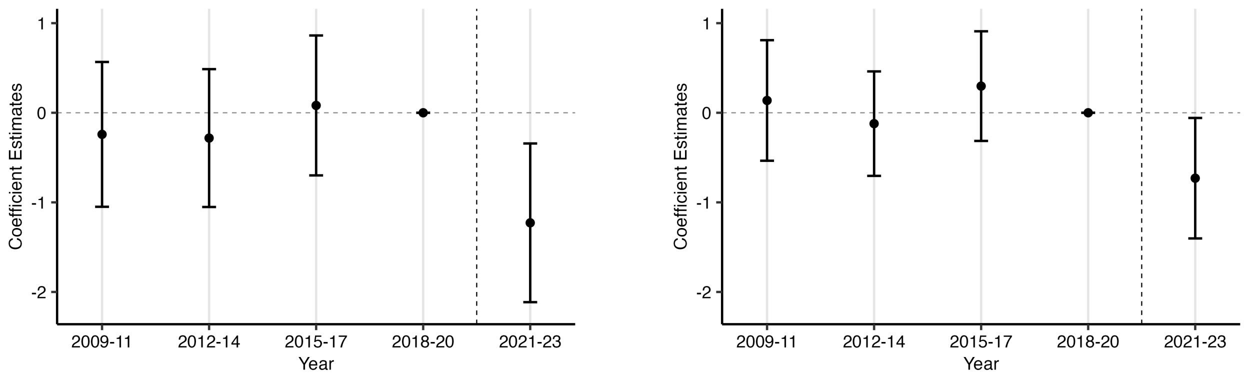
Figure 4 shows the estimation results. The post-policy logit coefficient on demolition permits is -1.23 in the boundary design, significant at the 1% level, and -0.73 in the matched design, significant at the 5% level, corresponding to decreases in the redevelopment rate of about 71 percent and 52 percent, respectively. The matched estimate is smaller in absolute value than the boundary estimate, though the difference is not statistically significant. The response is large relative to the burden the ordinances impose because they are temporary: a landlord can postpone redevelopment until they expire. Pre-policy demolition coefficients are insignificant in both designs, indicating parallel pre-trends in the treated and control areas. Construction permits fell by roughly as much as demolition permits in both designs, reflecting a reduction in redevelopment. The effect on renovation permits is not distinguishable from zero in either design (Figures D.4 and D.5). We additionally report the policy effects on rents, house prices, and permit processing times and costs, as well as estimates using alternative buffer widths, in Appendix D.

We discipline σ_c by matching the estimated demolition response to the two ordinances, both of which reduce redevelopment more when σ_c is larger.[34] Specifically, we impose the two anti-redevelopment policies for three years in the model on the set of parcels, within the neighborhood groups that the 606 Trail and Pilsen are in, that matches the observed pre-

[32]We match on pre-policy block-group level population density, median rent, home value, household income, build age, as well as demolition, construction, and renovation permit rates (Table D.2).

[33]We find that demolition and construction rates just outside the boundary did not rise after the policy (Figure D.3); rents and sale prices did not respond on either side (Figure D.8); and contractors previously active in the treated areas relocated to distant neighborhoods or stayed inactive rather than moving across the boundary (Table D.3).

[34]The surcharge lowers the log odds of redevelopment by $\sigma_c T$; the anti-deconversion ordinance removes the unit-reducing options, which are the profitable ones as redevelopment on average reduces the units (Figure 2). Proposition 3 formally discusses how each policy affects the redevelopment rate.



(a) Boundary design (b) Matched design

Figure 4: Difference-in-Differences Estimates for Demolition Permits

Note: The figure shows the logit estimation results of equation (16) for demolition permits, estimated on the boundary sample in Panel (a) and on the matched sample in Panel (b), both specified as described in the text. Conley spatial standard errors with a 100-meter cutoff. The reference period is 2018–2020, and 95% confidence intervals are displayed.

policy redevelopment rates and building characteristics of the policy areas. We solve for the landlord's problem on each treated parcel, holding rents fixed, and set σ_c so that the reduction in the log odds of demolition matches a coefficient of -1, which is between the two designs' estimates. Intuitively, we identify σ_c from the intertemporal decision that the landlords face: during the policy period, a landlord can either redevelop and be subject to the teardown tax and limit on the units or wait and redevelop once the policy expires. The observed fall in demolition therefore measures the extent to which landlords postpone redevelopment per dollar of policy-induced loss in the redevelopment value. We further assume landlords expected the ordinances to expire after three years as announced. Had they anticipated the extension that in fact followed, redeveloping later would have been affected as well, the observed decline in demolition would understate the response to a dollar of surcharge, and we would underestimate σ_c. See Appendix C for more details on the implementation.

We use the σ_c estimated from a three-year policy in two neighborhoods in a ten-year-period model of the entire city. The estimate carries over to the ten-year period because σ_c, the semi-elasticity of the redevelopment odds to a dollar of payoff, does not depend on the period length; however, the estimated construction costs do, which is analogous to fixed moving cost estimates in dynamic migration models (Howard and Shao, 2025). We therefore estimate σ_c at a three-year period, adjusting the discount factor, the depreciation rate, and the baseline redevelopment rates to the policy window, and then use the estimated σ_c to calibrate the costs at the ten-year period to match decadal redevelopment rates. We also assume that σ_c is common across neighborhoods, as is standard in models of housing supply

(Murphy, 2018). Consistent with this assumption, the policy effects estimated separately for the 606 Trail and Pilsen are similar (Figure D.6). We note that a common σ_c can still yield different effects of the same policy across neighborhoods in the model, since the same change in log odds moves the rate least where the baseline rate is near zero.

Construction and Renovation Costs

We calibrate the construction and renovation cost parameters $\{F_{\hat{q}x}, \Omega_x(h), C_x^{RN}\}_x$ and the number of parcels by q and h of each neighborhood. For ease of computation, we set $\mathcal{H} = \{0, 1, 2, 3, 4\}$.[35] To match the teardown rates of single-family and multifamily structures, which is important for studying the effect of the anti-deconversion ordinance, we introduce an added cost φ_x for demolishing multifamily structures. We assume the renovation cost function $C_x^{RN}(q', h)$ is a fraction c_x of the cost of building a new structure at the post-renovation quality $q' = (1 + \varrho)q$. We normalize the blueprint distribution $G(\hat{q})$ to be uniform over the quality grid, since it is not separately identified from the fixed cost schedule: more high-quality construction can reflect either more frequent high-quality draws or a lower cost of building at high quality.

We calibrate these parameters by neighborhood to match four sets of empirical moments on housing development inspired by the motivating evidence of Section 3. To compute the model moments, we solve each landlord's problem for one period in the model under myopic expectations in that they expect rents to stay at their estimated values and choose optimally given their parcel's state in 2020.[36] The first set of empirical moments is the *average redevelopment rate for single and multifamily housing.* The average level of fixed costs across blueprints governs the single family rate by reducing the profitability of redevelopment, while φ_x governs the multifamily rate. The second is the *difference in the quality distribution of new versus old buildings.* We compute the housing quality distribution separately for recently constructed and older structures, where we define new buildings as those built after 2005. This set of moments identifies how fixed costs vary across quality levels: if more new buildings are dominated by high quality structures in the model than the data, then the fixed costs to build high quality structures should be adjusted upward.

The third is the *difference in the housing unit-density distribution of new versus old*

[35] We make this assumption to reduce model dimensionality at minimal cost. Units in structures with 5–6 housing units comprise less than 5% of all units in structures between 1–6 housing units.

[36] Two features of our calibration are worth explaining. First, we do not impose that the baseline equilibrium is a steady state, nor are the data consistent with one. Second, we assume that landlords are myopic and expect current prices to persist. This is because sticking to rational expectations would require solving for the entire path of future prices along the transition. Transitions in the model are slow, so myopic land values stay close to their rational-expectations counterparts. We solve the perfect foresight equilibrium, in particular the entire sequence of the rent function, when conducting the counterfactual analysis Section 6. Appendix C.3 validates the approach.

Table 3: Summary of Model Parameters

Parameter	Description	Value	Source or target moment
Panel A: Externally calibrated parameters			
β	Discount factor (annual)	0.94	Average price-to-rent ratio
σ_x	Migration elasticity	6.3	Baum-Snow and Han (2024)
$\tau(x';x)$	Moving cost shifter, $x' \neq x$	0.8	Koşar et al. (2022)
η	Amenity elasticity to income	0.26	Macek (2024)
$\bar{L}(z, x_0)$	Income distribution and population	–	ACS
Panel B: Estimated and internally calibrated parameters			
δ	Quality depreciation rate (annual)	0.34%	Hedonic regression (15)
ϱ	Renovation quality premium	0.26	Hedonic regression (15), $\exp(0.23) - 1$
α	Housing preference weight	0.12	Expenditure shares by income group
$\bar{q}$	Stone–Geary minimum quality	1.07	Expenditure shares by income group
σ_c	Redevelopment elasticity (per $1,000)	0.033	606–Pilsen policy effect on demolition
Panel C: Neighborhood-specific parameters			
$F_{\hat{q}x}$	Fixed construction costs	–	Redevelopment rate (SF), New vs. old quality distribution
φ_x	Added cost of multifamily demolition	–	Redevelopment rate (MF)
$\Omega_x(h)$	Variable construction costs	–	New vs. old unit density
c_x	Renovation cost	–	Renovation rate
$\bar{A}(x, z)$	Exogenous amenities	–	Neighborhood income distributions

Note: Panel A parameters are calibrated externally. The migration elasticity is set below the 8.5 of Baum-Snow and Han (2024) to account for non-homothetic demand. Panel B and C parameters are estimated or calibrated internally, with the identifying moment listed. Panel C parameters vary by neighborhood; the construction cost estimates are reported in Figure C.2. SF and MF denote single-family and multifamily; the blueprint distribution is set uniform (Section 5.2.3).

buildings. This identifies the variable cost parameters $\Omega_x(h)$ by unit: if newly developed structures contain more single-family homes than the existing stock, the incremental cost of building multi-family structures must be high. We normalize $\Omega_x(1) = 0$, so the fixed costs is the cost to build a single family home of quality $\hat{q}$. In Figure C.1, we show that new buildings on average have 20% fewer units and 50% higher quality relative to older buildings. The fourth is the *renovation rate*, which identifies the renovation cost factor c_x. Finally, we set the number of parcels over (q, h) states to match the quality distribution $H(q, x)$ recovered from the demand block.

5.3 Estimation Results

Table 3 summarizes the model parameters and the source or moment that identifies each. For the externally calibrated parameters, we set the annual discount factor to 0.94 to match the average price-to-rent ratio in Chicago. Following Baum-Snow and Han (2024), we set the migration elasticity to $\sigma_x = 6.3$, which targets a migration elasticity to wages of 8.5 that their estimates imply. Following Koşar et al. (2022), who estimate renters' median willingness to pay for a local move at 22.45% of income, we set $\tau(x'; x) = 0.8$ for all $x \neq x' \in \mathbb{X}$. Using the plausibly causal estimates of Macek (2024), we set the elasticity of amenity values to income at $\eta = 0.26$, and invert the exogenous amenities $\bar{A}(x, z)$ from the observed income distributions.[37] We obtain the spatial income distribution and population from the ACS for the entire Chicago–Naperville–Elgin, IL-IN-WI MSA and set the areas outside the municipality as the outside option.

On the housing demand parameters, we calibrate the housing demand parameters to be $\alpha = 0.12$ and $\bar{q} = 1.07$. These imply that the highest-income households spend about 12% of their expenditure on housing, and the Stone–Geary floor raises the share to roughly one-half at the bottom of the income distribution. Evaluated at the city-average rent function, the value of $\bar{q}$ implies that the lowest income household will spend at least \$500 per month on rent. With only two parameters, the model tracks the full expenditure share profile closely, as in Couture et al. (2024).

On the cost parameters, we calibrate the redevelopment elasticity to be $\sigma_c = 0.033$ per thousand dollars of payoff. At the calibrated σ_c, the surcharge and the anti-deconversion ordinance each account for roughly half of the model's decline in redevelopment.[38] At the citywide average demolition rate of roughly 2 percent per decade, an increase in the redevelopment payoff equal to 1 percent of the median home value in Chicago—roughly \$3,000—raises the demolition rate by about 0.2 percentage points per decade. Our calibration procedure also yields reasonable estimates of construction costs (Figure C.2). We estimate that building a single family home costs \$120,000 at the 10th percentile of quality, increasing to \$175,000 at the median quality and \$700,000 at the 98th, with the most expensive neighborhood–blueprint pairs exceeding \$1.4 million.[39] Costs vary nearly as much across neighborhoods as

[37] Macek (2024) estimates that households are willing to pay approximately 0.23% of their income to live in a neighborhood with 1% greater average income. Our chosen value of 0.26 targets this estimate, accounting for differences in the marginal utility per dollar in our demand model.

[38] Because the ordinance binds only on multifamily structures, the model predicts a larger demolition response for multifamily than for single-family parcels. Estimated separately by structure type, the two treatment effects are not significantly different, but the treated areas contain too few parcels of either type to detect a gap of the size the model implies.

[39] EcoBuild Plus is a Chicago area contractor that specializes in the construction of single family and low-density multifamily homes. They estimate that the total costs of building an average home range from

across quality segments: fixed costs are higher and more dispersed in higher-income neighborhoods, and the extra cost of multi-family relative to single-family construction is larger in these neighborhoods. This likely reflects regulatory restrictions on housing density in affluent neighborhoods (Macek, 2024). On renovation, the calibrated renovation cost parameter c_x has a median of 0.93 across neighborhoods, so a major renovation costs roughly as much as constructing the target quality anew.

5.4 Model Fit

We assess model fit on three sets of untargeted moments across neighborhoods. The first is average housing quality. We recover each neighborhood's quality distribution from its income distribution and the estimated rent function through the household's utility function, rather than using the hedonic estimates. Model quality tracks average hedonic quality across neighborhood groups (Figure C.3), which validates our housing demand specification. The second, average rent, is a by-product of the first: with the rent function estimated from the data, model and empirical average rents differ only through the quality distribution, and the two align closely (Figure C.4). The third is the price-to-rent ratio. In the model, a higher price-to-rent ratio is associated with a greater share of the redevelopment option value in the home value, which governs the redevelopment rate. Model ratios are tightly correlated with the data (Figure C.5); however, the ratio is far less dispersed in the model as it abstracts from many determinants of housing prices, including expected rent and income growth, rent risk, and credit conditions.

6 Counterfactual Analysis

We use the estimated model to evaluate housing policies in general equilibrium. We begin with an expansion of the anti-deconversion ordinance and the teardown tax policy bundle. We then compare two place-based alternatives with the same goal of improving housing affordability for low-income households: a redevelopment subsidy in high-income neighborhoods and a rent subsidy in low-income neighborhoods. We evaluate each policy by solving for the full transition path and comparing it to that of the baseline economy without the policy.

\$450,000 to \$700,000 inclusive of land acquisition costs. For luxury homes, costs can easily exceed millions, with prices up to \$500 per square foot. These costs are inclusive of additional fixed costs for architecture and interior design consultation. See their article here.

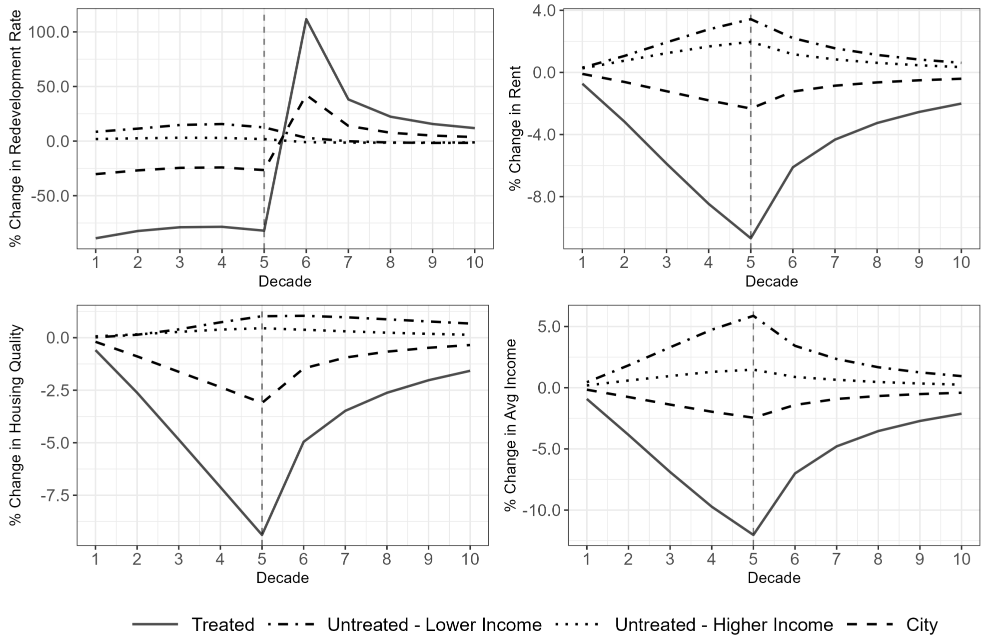


Figure 5: The Impact of the Anti-Redevelopment Policy on Housing Redevelopment, Rent, Quality, and Neighborhood Income

Note: Percent change relative to the baseline economy in the same decade. The vertical line marks the last decade in which the policy is in force.

6.1 The Anti-Deconversion Ordinance and the Teardown Tax

Chicago is scaling up the two policies studied in Section 5. Before the 606–Pilsen demolition surcharge expired, the City Council approved the Northwest Side Preservation Ordinance, which extends the anti-deconversion ordinance to more neighborhoods and raises the surcharge fourfold, to $60,000, through 2030.[40] We simulate an expansion of this form, imposed for 50 years, on neighborhoods with below-median income and above-median redevelopment rates. These neighborhoods are the most exposed to gentrification.

Average neighborhood and citywide outcomes. Figure 5 reports the policy effects on the redevelopment rate, average rent, average housing quality, and average income separately for the treated neighborhoods, untreated neighborhoods split by average income, and the city as a whole; Figure F.7 maps the three neighborhood groups. In the treated neighborhoods, the decadal redevelopment rate falls by 82 percent while the policy is in force. With the decrease in redevelopment, the housing stock depreciates, and the quality distribution shifts downward. Landlords substitute toward renovation, but the response is inadequate to offset the quality decline (Figure F.8c). By the last decade of the policy, the treated housing stock

[40]See the City of Chicago's policy description here.

is 9.4 percent lower in average quality, its residents are 12 percent poorer, and its average rent is 10.7 percent lower than in the baseline. The policy therefore achieves its goal in the treated neighborhoods: redevelopment falls sharply and housing becomes more affordable. Average rent falls for two reasons. The housing stock is of lower quality, and the neighborhood is poorer, which lowers housing demand and shifts the rent function down. The resulting decline in amenities lowers demand further, as Figure 3 illustrates.

Untreated neighborhoods are also affected by the policy through general equilibrium effects. The relatively high-income households that leave the treated neighborhoods migrate to the untreated ones, and their migration increases housing demand within the untreated neighborhoods, especially for high-quality housing, which raises average income, rent, and the return to redevelopment there. Untreated lower-income neighborhoods are affected more than untreated higher-income neighborhoods. This is because the arriving high-income households raise average income by more in the untreated low-income neighborhoods, where it rises by 5.8 percent, about four times the increase in the untreated high-income ones. The resulting response in neighborhood amenities leads to greater increases in housing demand in untreated, low-income neighborhoods. Therefore, the policy reduces redevelopment and gentrification in the treated neighborhoods but reallocates them toward the untreated neighborhoods that most resemble them.

At the city level, the policy lowers average rent by 2.3 percent at the end of the policy period, while also reducing housing quality and average income by 3.1 and 2.3 percent respectively. The policy makes the city more affordable yet less attractive at once, resulting in net out-migration of high-income households. Once the policy expires, redevelopment in the treated neighborhoods overshoots its baseline in the following decade, as landlords carry out the projects they postponed to avoid the policy restrictions, and the other outcomes recover slowly toward baseline.

The rent function and quality distribution. Figure 6 reports the change in rent and in the number of units at every quality level at the end of the policy period. These changes matter for heterogeneous welfare consequences among households because households with different incomes consume different qualities. Holding amenities fixed, a household is better off when the same quality unit rents for less.

The left panel reports the change in the number of housing units along the quality distribution, as well as the change in the total number of units. In the treated neighborhoods, the anti-redevelopment policy preserves low-quality housing but reduces the supply of high-quality housing, cutting the number of units at the top of the distribution by about a third. Contrary to the policy's intent, the total number of habitable units there falls: although

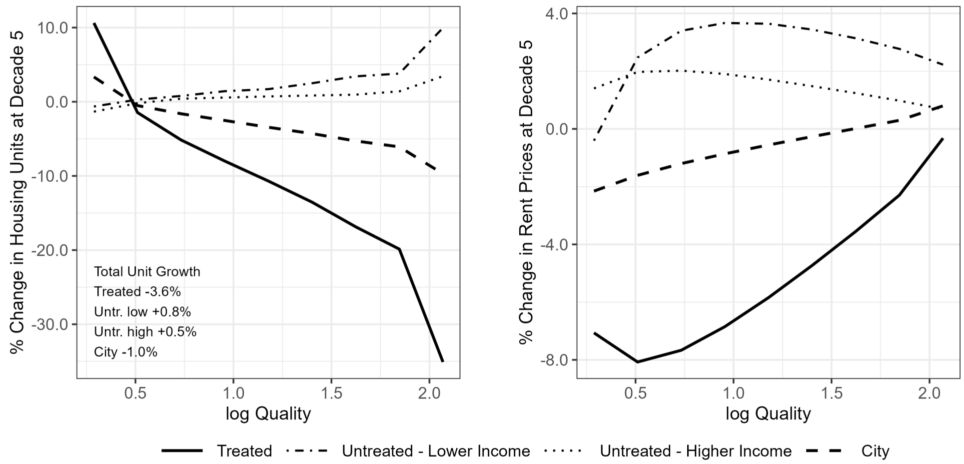


Figure 6: The Impact of the Anti-Redevelopment Policy on Units and Rent by Quality

Note: Percent change in the number of housing units (left) and in the rent function (right) against log housing quality, in decade 5, relative to the baseline economy. The inset reports the percent change in the total number of units for each neighborhood group. We exclude from the unit counts units that have depreciated to the minimum quality $\bar{q}$ and are no longer habitable.

the anti-redevelopment policy is meant to preserve affordable units, many old housing structures depreciate until they are no longer livable and are not redeveloped.[41] In the untreated neighborhoods, the increase in redevelopment replaces low-quality housing with high-quality housing. The increase in housing demand also leads to an increase in the number of housing units supplied via redevelopment and renovation. Citywide, the quality distribution shifts toward low-quality housing, and the number of habitable units falls by about 1 percent.

The right panel reports changes to the rent function. In the treated neighborhoods, rents fall throughout the quality distribution, and by more for low-quality housing. The downward shift in quality steepens the rent function, as Proposition 2 predicts. In the untreated neighborhoods, rents rise throughout the quality distribution, except for the lowest-quality housing. Interestingly, the rent for medium quality housing increases the most in untreated neighborhoods, and more so in the untreated low-income ones. This is explained by two forces. First, migrants to untreated neighborhoods tend to be of medium and high income, which puts upward price pressure on medium and high quality segments. Second, increasing housing demand in untreated neighborhoods causes households to substitute down the quality ladder, putting more price pressure on medium rather than high quality housing; this is the "trickle-down" effect in reverse (Nathanson, 2025). Citywide, rents fall for low-quality housing and rise for high-quality housing.

[41] In a model where smaller renovations are possible, we acknowledge that units could be preserved as landowners substitute away from demolition in response to each policy.

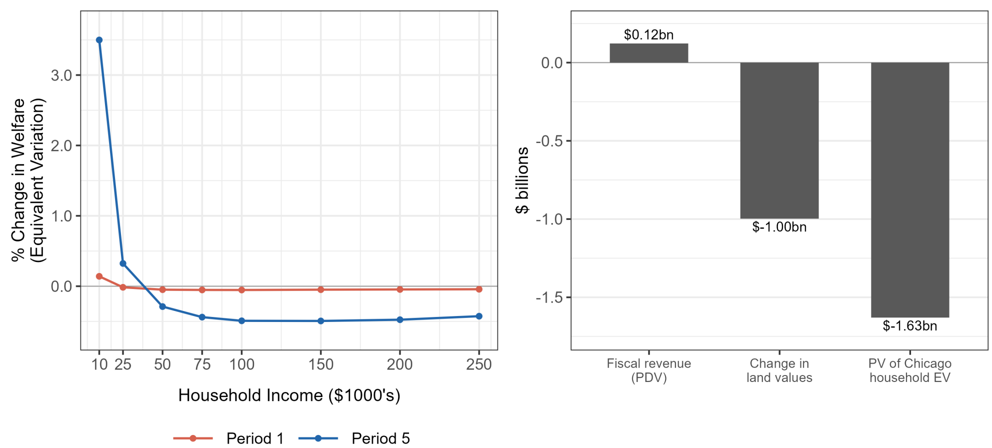


Figure 7: The Impact of the Anti-Redevelopment Policy on Welfare and City Aggregates

Note: The left panel reports the equivalent variation as a share of income, by household income, in the first and last decade of the policy. The right panel reports the changes in the discounted values of fiscal revenue and of household equivalent variation over the model horizon, as well as the change in land values in period 1 of the policy, which captures how the policy affects future rents and redevelopment opportunities. The present discounted value of household equivalent variation is calculated only for households whose origin is within the Chicago municipality.

Welfare. Figure 7 reports the welfare effects by household income and the city-aggregate changes in the present value of fiscal revenue, land values, and household welfare.[42] The left panel shows the welfare changes of households with different income. In each period, we measure the welfare effect on a household as the equivalent variation, which is the percentage change in income that would leave households as well off as under the policy (see Appendix E.1 for details). The policy redistributes welfare toward the poorest households and away from everyone else: at the end of the policy period, households earning $10,000 gain 3.5 percent of income, while all households with income above $35,000 lose. Households making $250,000 lose by 0.43 percent, which is less than the 0.50 percent lost by households earning $100,000. Highest-income households spend the smallest share of their income on housing and are therefore less sensitive to changes in rent. The presence of moving costs implies that low-income households initially living in the treated neighborhoods gain somewhat more than those initially living elsewhere (Figure F.9).

The right panel shows that the policy costs $1.63 billion in present discounted value of household welfare, even though average rent falls citywide. Welfare falls for all but the lowest-income households because the rent decline reflects two forces that do not make households better off. One is quality downgrading: average rent falls partly because households occupy lower-quality units. By contrast, in standard models with divisible housing services, every

[42] Throughout the counterfactual analysis, we compute welfare changes only for households initially residing in Chicago.

household faces one price per unit of housing services, so a lower price benefits every renter when amenities are held fixed; under non-homothetic preferences, low-income renters gain more (Gaubert and Robert-Nicoud, 2025; Finlay and Williams, 2025). With indivisible housing, the rent schedule is nonlinear and need not shift uniformly, so welfare changes can be positive for some income groups and negative for others. The other force is a compensating differential: as neighborhood amenities deteriorate, rent must fall to keep households willing to live in a worse neighborhood.

Land values fall by \$1.00 billion across the city in period 1. The loss is borne by the landlords in treated neighborhoods, so a policy intended to protect low-income renters transfers wealth away from landlords in the neighborhoods it targets. On net, the policy achieves the goal of improving housing affordability for the low-income households but is costly: the \$0.12 billion in present discounted tax revenue collected from the demolition tax is far smaller than the combined losses of households and land value.

The role of endogenous amenities. Endogenous amenities significantly amplify the policy's effects on the treated neighborhoods and its spillovers to the untreated ones. To quantify the importance of this channel, we re-solve the model holding amenities fixed at their baseline values and report the results in Figure F.10. In the treated neighborhoods, the effects on average rent and average income shrink by more than half, while the declines in redevelopment and housing quality are similar. In contrast, the spillover effects to untreated neighborhoods are much smaller than their endogenous-amenity counterparts; this is driven by much less migration of high-income households in the absence of self-reinforcing amenity changes.

The welfare losses from the policy are also much smaller when amenities are exogenous, as shown in Figure F.12. Under exogenous amenities, welfare changes are positive for households earning up to $\$75,000$ (instead of $\$35,000$). Summed over all households, the policy costs \$0.30 billion in present discounted value of household welfare, less than a fifth of the \$1.63 billion loss under endogenous amenities, and land values fall by \$0.37 billion rather than \$1.00 billion. With fixed amenities, rent need not fall to compensate households for a worse place to live, and the loss that remains reflects quality downgrading alone.

The effect of the anti-deconversion ordinance. Our analysis indicates that the bundle of the two policies imposes serious costs. Could the anti-deconversion ordinance alone achieve the same goal at a lower welfare cost? Imposed alone, the ordinance cuts redevelopment in the treated neighborhoods by roughly half as much as the bundle, yet lowers rent and average income there by nearly as much, with a smaller decline in housing quality (Figure F.13). Compared to the policy bundle, it also preserves units: the treated housing stock is

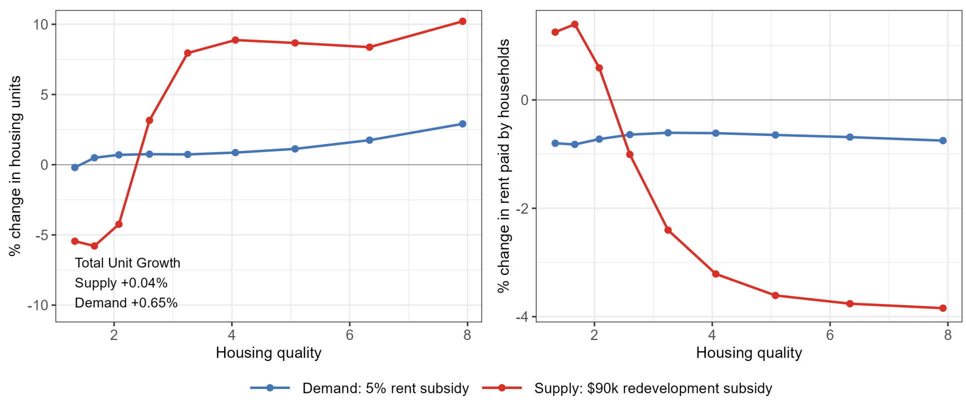


Figure 8: The Impact of the Redevelopment and Rent Subsidies on Units and Rents by Quality

Note: Percent change in the number of housing units (left) and in the rent paid by households (right) against housing quality, citywide, in decade 5. Rent is measured net of the rent subsidy where it applies. The inset reports the percent change in the total number of units under each subsidy relative to the baseline economy in Decade 5.

0.8 percent larger than in the baseline, against 3.6 percent smaller under the bundle (Figure F.14). The ordinance alone nonetheless produces a sizable welfare loss, $0.98 billion in present discounted value, about half the $1.61 billion under the bundle and in line with its effect on redevelopment. The ordinance preserves too few units to offset much of the welfare loss.

6.2 Assessing Alternative Policies to Improve Housing Affordability

We evaluate two additional policies aimed at improving housing affordability for low-income households: a redevelopment subsidy of $90,000 per project in above-median-income neighborhoods, which expands new housing supply where housing is expensive, and a rent subsidy covering 5 percent of rent for households in below-median-income neighborhoods, which eases their rent burden directly. To make the two comparable, both last 50 years and cost almost exactly the same.

In a simple supply-and-demand model of housing, the choice between a subsidy to supply or demand does not matter for economic incidence. Subsidies of the same fiscal size move price and quantity by the same amount, with the relative elasticities of supply and demand determining how the benefit is split between renters and landowners. By contrast, our framework features vertically differentiated housing, redevelopment that increases the average quality of the housing stock and need not expand the supply of units, and amenities that respond to who lives where. Each of these model features will cause the supply and demand subsidies to differ by incidence and deadweight loss.

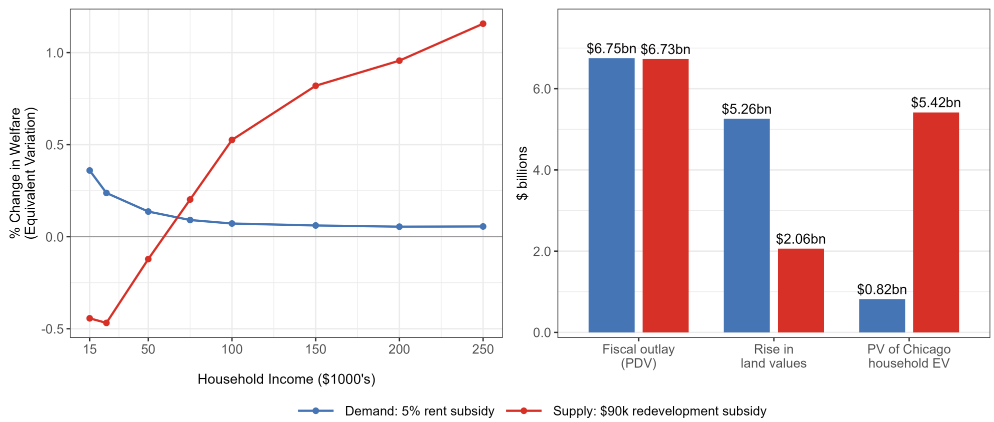


Figure 9: Welfare and Fiscal Incidence of the Redevelopment and Rent Subsidies

Note: The left panel reports the equivalent variation as a share of income in decade 5, by household income. The right panel reports the period-1 present discounted values of the fiscal outlay over the policy window, the change in aggregate land values, and household equivalent variation over the model horizon. The present discounted value of household equivalent variation is calculated only for households whose origin is within the Chicago municipality.

The rent function and quality distribution. Figure 8 reports each policy's effect on the number of housing units along the quality distribution, together with the total number of units, and on the rent paid at each quality at the end of the policy period. The redevelopment subsidy upgrades the housing stock: units shift from the bottom of the quality distribution to the top, while the total number of units is essentially unchanged. Rents move with the stock, falling by about 4 percent for high-quality housing and rising for low-quality housing. The subsidy therefore fails to deliver the trickle-down benefits to low-income households. Most of the new high-quality units are filled by high-income households migrating into the subsidized neighborhoods, many from outside the city, so few low-quality units are vacated, consistent with the limited trickle-down of high-quality housing supply that Nathanson (2025) attributes to in-migration. Moreover, unlike greenfield development, redevelopment reduces the low-quality housing the poor live in, raising their housing costs even further.

The place-based rent subsidy targets the low-income population more directly because it pays households where they already live. It decreases the out-of-pocket rent expenditure almost uniformly across households. Most of the subsidy, however, is captured by landlords through the large capitalization of the subsidy into rents: on average, only around 1.25% of the rent subsidy is passed through to the households (Figure F.15). Because rents rise along the quality distribution, landlords who redevelop under the rent subsidy build more units per project, which expands the housing stock modestly, by about 0.7 percent.

Welfare. Figure 9 reports the welfare effects by household income and the city-aggregate changes in fiscal outlay, land value, and household welfare under each subsidy. The redevelopment subsidy is regressive: households below about $60,000 lose, while households at $250,000 gain the equivalent of 1.15 percent of income. Under the rent subsidy, every household gains, and the gains decline with income, tracking the rent share of income. Each subsidy benefits the households who consume the segment of the stock it makes cheaper. Summed over all households, the redevelopment subsidy delivers $5.42 billion in present discounted household welfare, against $0.82 billion under the rent subsidy. Much of the difference comes from the amenity gains brought by the redevelopment subsidy, which we discuss later.

The two subsidies also differ in their effects on total land value. The redevelopment subsidy raises the option value of redevelopment but lowers the city's average rent, so land values rise only modestly. The rent subsidy increases the land value by more, as 78% of the subsidy is absorbed by higher rents and capitalized into land values. Per dollar spent, the redevelopment subsidy returns $1.11 in combined household and landowner benefits, while the rent subsidy returns only $0.90 to Chicago households and landowners combined.

The role of endogenous amenities. Figure F.16 reports the city-wide policy effects when we assume exogenous amenities. Comparing with the results in the baseline model shows that the redevelopment subsidy's large welfare gain comes from the endogenous amenities channel. Thus subsidy draws more high-income households into the richest neighborhoods, further raising amenities there. Quantitatively, the total household welfare gain decreases by about 60% without endogenous amenities. By contrast, the rent subsidy induces relatively less resorting of households with different income across neighborhoods, so endogenous amenities play a small role in its welfare effects.

6.3 Implications for Housing Policies

Although a full analysis of optimal housing policies is beyond the scope of this paper, we discuss what welfare-improving policies would look like through the lens of our model.[43] Inefficiency arises in the model because households do not internalize how their neighborhood choices affect the amenities of incumbent residents. In particular, middle- and low-income households free ride on the amenities created in rich neighborhoods, lowering the average income and the amenity value that attracted them in the first place. The problem is analogous to the fiscal externalities that arise with the provision of local, congested public goods

[43]Solving a planner's problem for the city would require taking a stand on welfare weights and on households outside the municipality, and we stop short of it.

(Calabrese et al., 2012).[44] Since the housing quality distribution significantly affects a neighborhood's income distribution, the redevelopment decisions of landlords do not internalize effects on neighborhood amenities. In addition, under non-homothetic preferences, low-income households benefit more from cheaper housing and are willing to trade off lower amenities. A welfare-improving policy should therefore combine two margins, promoting greater income sorting for high-income households while providing affordable housing in low-income neighborhoods. This is consistent with efficient Tiebout sorting (Calabrese et al., 2012).[45]

The mitigation of free riding explains why the redevelopment subsidy generates the largest gain in household welfare: by letting high-income neighborhoods absorb more high-income households, it prices out the free riders and brings the sorting of households across neighborhoods closer to what a planner would choose. The anti-redevelopment policy seemingly delivers both margins, since it results in cheaper housing in low-income neighborhoods and greater income sorting. However, it also lowers the city's average household income, distorts housing markets by reducing housing quality, and shrinks the stock of habitable housing. This leads to massive deadweight loss.

7 Conclusion

Redevelopment is a major source of new housing in dense cities, yet its welfare consequences are hard to assess: it upgrades quality more than it adds units, it clusters in particular neighborhoods, and because housing is durable, its effects unfold over decades. We build a dynamic spatial general equilibrium model that captures all three features. Within each neighborhood, an assignment model matches households of different incomes to units of different quality and delivers a rent function over quality. Across neighborhoods, households sort by income, and amenities respond to who lives where. Over time, the stock depreciates and forward-looking landlords choose when to redevelop. Rents, sorting, and redevelopment are thus determined jointly. We discipline the redevelopment elasticity with difference-in-differences evidence from anti-redevelopment policies in Chicago.

We apply the model to assess the scaled-up anti-redevelopment policy. The policy achieves its goal where it applies, reducing redevelopment and gentrification and making rental housing more affordable in the treated neighborhoods. It also reallocates both toward the untreated neighborhoods that most resemble them, lowers the city's average income and stock of habitable housing, and benefits only the poorest households, with private losses far exceeding the

[44] Calabrese, Epple and Romano (2007) study the role of housing policy in correcting this externality in the context of minimum-lot-size restrictions; Macek (2024) does so in a quantitative spatial model.

[45] Because relocation is costly, the gains from any place-based policy accrue mostly to the households already living where it applies, and a planner who reallocates households across neighborhoods must also count the moving costs the reallocation induces.

tax revenue it raises. Among alternative policies with the same budget, a place-based rent subsidy improves affordability for the poor and roughly returns what it costs, though most of it is capitalized into land values, while a redevelopment subsidy in high-income neighborhoods nearly doubles its budget in total benefits through the endogenous-amenity channel, yet is regressive and generates no trickle-down effect to low-quality housing.

One important takeaway of the quantitative exercises is that in a housing market differentiated by quality and neighborhood amenities, affordability measured by average rent is a poor guide to welfare, and policies that appear equivalent in a single-market framework diverge sharply. Beyond the policies we examine, the framework is well suited to study policies that change local housing quality distributions, such as the Low-Income Housing Tax Credit and the demolition of public housing.

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Appendix

A Data Appendix

A.1 Parcel Definitions and Permit Matching

This appendix documents how we define the parcel—the unit of observation in all of our empirical work—and how we assign building permits and assessment records to it. Section 2 summarizes the procedure; here we give the construction in full.

A.1.1 The fixed 2005 parcel map

A *parcel* is a single legally defined lot of land. Cook County publishes a map of every parcel in the county, in which each lot is a polygon carrying a fourteen-digit property identification number (PIN). The first ten digits of that number identify the lot itself; the last four distinguish individual condominium units within a building standing on it. We work at the level of the ten-digit identifier, which we refer to as the parcel.

Lots do not stay fixed. Owners divide one lot into several, combine several into one, and the county periodically re-surveys boundaries. We therefore fix the map at a single vintage—the 2005 edition—and hold it constant for the entire sample period. We take the county-wide 2005 map, retain the polygons that fall within the City of Chicago, and keep one polygon per ten-digit identifier. This yields **602,839 parcels**, each of which receives a permanent identifier used throughout the analysis. Because condominium units in the same building are recorded as separate rows sharing essentially the same polygon, keeping one row per ten-digit identifier collapses them to the single lot they occupy.[46]

We delimit the city using the union of Chicago's community-area boundaries, which is the city-boundary definition used elsewhere in our data construction. We apply the identical boundary to the 2024 map described in Section A.1.4, so that the two vintages are cut in the same place and cannot disagree at the city's edge merely because two different boundary files were used.

A.1.2 Locating assessment records on the map

Assessment records are reported for a PIN and a year, and the assessor also publishes a geographic coordinate for each PIN and year. We use those coordinates to place every assessed property on the 2005 map, by testing which polygon the point falls inside.

[46] We retain one representative polygon rather than merging the duplicate polygons geometrically. The two are equivalent for determining which lot a point falls on, which is the only use we make of the polygons; they would differ for calculations of lot area, which we do not perform.

A PIN's recorded coordinate moves slightly from year to year, by well under one meter, because the assessor re-geocodes addresses periodically. These displacements are far smaller than a city lot, so every year of a given PIN would resolve to the same polygon. We use this directly: rather than testing one point per PIN-year, we take each PIN's most recent valid coordinate, test it once, and carry the resulting parcel identifier to every year of that PIN's history. Besides being faster, this guarantees that one physical property is attached to the same parcel for its entire recorded history, which testing year by year would not. Of the **733,089** distinct Chicago PINs with a usable coordinate, **99.95%** fall inside a 2005 parcel.

We then parse each assessment address into a house number, a directional prefix, a street name, and a street-type suffix—for example, "1200", "N", "ASHLAND", "AVE"—so that it can be compared with the addresses recorded on permits. Two details make this comparison reliable. First, we build the list of admissible street-type suffixes from the permit records themselves, so that both sides of the match draw on the same vocabulary rather than on two independently compiled lists. Second, when a suffix appears more than once in a string we take the last occurrence, so that a street whose name happens to contain a word such as "PARK" is not truncated at the wrong point; anything following the suffix, such as a condominium unit number, is discarded. We apply the same handful of street-name corrections—for spellings such as "MCLEAN" versus "MC LEAN"—to both the assessment and the permit side. Finally we collapse the records to one observation per address and year, since condominium units in one building share an address once the unit number is removed.

A.1.3 Matching permits to parcels

Permits record an address but no parcel identifier, so they must be matched. We proceed in two steps.

Step one: match through the assessment record. We look for an assessment record at the permit's address and take the assessment year closest to the permit. Which side we search depends on the permit type, because the question each permit type poses is different. For demolition and renovation permits we take *the most recent assessment at or before the year the permit was issued*: what we want to know is what stood on the lot before the permitted work, since that is what is being torn down or altered. For new construction we instead take *the earliest assessment at or after the permit year*. New construction is built at addresses that often have no assessment history at all until the finished building is reported to the assessor, which lags the permit; searching backward would discard these permits, while searching forward finds the first record that reflects the completed building.[47] We match first

[47]The asymmetry is visible in the data: 78.5% of demolition addresses have an assessment record predating their permit, against 41.1% of new-construction addresses. A third of new-construction addresses have their

on house number, direction, street name and suffix; for permits still unmatched we repeat the match without the directional prefix. Whenever a permit matches several assessment records, the timing rule selects among them.

Step two: match through the building-footprint map. Permits whose address appears nowhere in the assessment roll are matched instead to the city's map of building footprints, using the same address fields. Where a footprint covers a range of house numbers we require the permit's number to lie inside the range; where no exact match exists we accept the nearest house number on the same street within one hundred. We then take the matched building's coordinate and test which 2005 parcel it falls inside.

Table A.1 reports the outcome. Overall, 98.5% of the 187,867 permits in our sample are assigned to a parcel, and the permits reach 101,485 distinct parcels. Every permit records which of the two routes produced its match, so that specifications can be restricted to the assessment-matched subset when the pre-permit characteristics of the building are themselves needed.

Table A.1: How permits are matched to 2005 parcels

	Assessment			
Permit type	Before permit	After permit	Building footprint	Unmatched
Demolition	79.2%	—	19.8%	0.9%
New construction	—	74.9%	24.1%	1.0%
Renovation	54.1%	—	44.2%	1.7%
All permits	51.8%	6.6%	40.1%	1.5%

Notes: Shares of permits by the route through which each was assigned to a 2005 parcel, as described in Section A.1.3. Demolition and renovation permits are matched to the most recent assessment at or before the permit year; new construction to the earliest assessment at or after it. Rows sum to 100%. The sample is the 187,867 demolition, new-construction and renovation permits issued between 2006 and 2023.

A.1.4 Parcel lineage and new construction

A lot that existed in 2005 may since have been divided or combined, so a building constructed today may sit on land that was one 2005 lot, part of one, or parts of several. To handle this we build a second parcel map from the 2024 county file, clipped to the city exactly as the 2005 map is, and compute the geographic overlap of every pair of 2005 and 2024 lots.

We record a pair as *materially overlapping* when the shared area is at least 20% of the area of the smaller of the two lots. A threshold is necessary because the two maps are drawn

first assessment record appear only after the permit, with a median lag of one year.

independently: adjoining lots routinely overlap slightly through nothing more than survey and digitization differences. The threshold is not arbitrary. The distribution of overlap shares is sharply bimodal rather than smooth: roughly 685,000 overlapping pairs share less than 15% of the smaller lot, roughly 593,000 share more than 70%, and only about 25,000 lie in between. Chicago's residential lots are narrow—about twenty-five feet wide—so an offset of a few feet between two independently drawn maps spills more than a tenth of a lot's area onto its neighbor. The 20% threshold sits inside the empty region between the two clusters, so that moving it moderately in either direction reclassifies few pairs.

Classifying each 2005 lot by what became of it, 53.6% are unchanged, 41.8% correspond one-to-one to a single 2024 lot with a redrawn boundary, 2.0% were absorbed into a larger lot, 0.5% were divided, 1.9% have both relationships, and 0.2% have no material overlap with any 2024 lot. Lots with a division or combination in their history—4.4% of the city—are the cases where the identity of the underlying land is genuinely ambiguous. We caution against reading the 41.8% category of redrawn boundaries as economic reorganization of land: it mixes genuine lot-line adjustments with differences in how the two maps were surveyed and drawn, and we have not validated the distinction.

For new-construction permits only, we then assign the permit to *every* 2005 lot that materially overlaps the present-day lot it was built on, rather than only to the single lot its coordinate falls inside. This is additive: the original match is always retained and flagged, and 2,783 additional permit-parcel pairs are created, so that the 187,867 permits generate 190,650 permit-parcel observations. Demolition and renovation permits are never expanded in this way. Because a construction permit can therefore appear against more than one parcel, any measure that counts permits must say how it treats the duplication. Our parcel-year panel records at most one permit of each type per parcel and year, so the expansion cannot double-count within it; a measure that instead counts permit records would need to restrict to the flagged primary match.

A.1.5 Coverage and unmatched permits

The 1.5% of permits (2,911) that reach no parcel divide into two groups. The large majority—2,584, or 89%—are renovation permits at addresses appearing in neither the assessment roll nor the building-footprint map. Both sources are inventories of residential and small commercial structures, so permits issued on airports, hospitals, schools, industrial sites and municipal property fall outside both by construction; O'Hare Airport is among the recurring addresses. This is a coverage limitation of the underlying sources rather than a failure of the matching procedure, and it is concentrated in a part of the building stock our analysis excludes in any case. The remainder had a footprint coordinate that fell inside no 2005 lot, which occurs mainly on land that was not platted as a distinct parcel in 2005. Unmatched permits are distributed across years roughly in proportion to permit volume, so they are not an artifact of the early part of the sample.

A.1.6 Unit counts

Two counts of dwelling units per parcel are used in the analysis, and they serve different purposes.

The first is a fixed count taken from the building-footprint map, which records the number of units in each structure. We place every footprint on the 2005 map and sum units across all structures standing on the same lot, since one lot may carry more than one building. This covers 472,630 parcels, or 78.4% of the city, of which 398,920 have between one and six units. This is the count we use to define the sample of one- to six-unit residential parcels, because it is available for the whole city and does not depend on a parcel ever having been permitted or assessed. Its limitation is that it is a single present-day snapshot: a lot rebuilt during the sample period is described by what stands on it now.

The second is a time-varying count taken from the assessment records, giving the assessed number of units for each parcel and year. Where several PINs map to the same 2005 lot—condominium units in one building, for instance—we take the most commonly reported value. This yields 11,076,755 parcel-year observations covering 456,868 parcels over 2000–2024. It is used where a parcel's status must be evaluated as of a particular year, by carrying forward the most recent assessment at or before that year, so that a single-family home later replaced by a multifamily building is not treated as multifamily in the years before its replacement.

A.1.7 Interpretation and limitations

Fixing the parcel map has a clear interpretive consequence. Our redevelopment measures capture the intensity of building activity on a constant base of land, not a count of distinct legal transactions or distinct development projects. When several 2005 lots are assembled and redeveloped together, each of the underlying lots is recorded as having been built on, which is the correct statement about land but overstates the number of separate decisions. When one lot is divided and built on twice, the reverse ambiguity arises. We prefer the fixed map nonetheless, for the reason given in Section 2: a map that moved with the assessor's contemporaneous definitions would let developers change the denominator of a redevelopment rate by dividing or combining lots, even with physical building activity held fixed. This is the same reason empirical work on neighborhood change holds census geography constant.

Two limitations follow and are worth stating plainly. First, our structural model treats a parcel as a fixed unit of land on which the owner chooses whether to rebuild, and does not model the assembly or subdivision of land as a separate decision. The empirical measure is aligned with that abstraction, but neither the model nor the measure can speak to policies that operate mainly by making it easier or harder to reorganize lots. Second, where an analysis compares a parcel's characteristics before and after redevelopment using a common identifier, lots that were divided or combined generally acquire new identifiers and so drop out of the comparison. Those results describe redevelopment on parcels whose identifiers persist, which need not be representative of redevelopment as a whole; assembly and subdivision are plausibly associated with larger changes in structure size than ordinary same-parcel replacement, so the surviving sample may understate the dispersion of structural change. We do not claim that the resulting measurement error is classical or that it necessarily attenuates any particular estimate.

A.2 Rent Imputation

This appendix details the imputation of market rents described in Section 5.1. We link RentHub listings to assessment records by address; 41 percent of RentHub addresses match, covering 29,562 parcels. Assessment records cover residential buildings with one to six units, so listings in larger apartment buildings have no counterpart and are dropped. Unit characteristics come from the assessment records rather than the listings. Because these records are kept at the building level, we assume units are homogeneous within a building and divide assessed value, bedrooms, bathrooms, and floor space equally across units.

On the linked sample, we regress the within-year rent rank of listed units on a quadratic in the within-year per-unit assessed-value rank, the characteristics in $\mathcal{C}$, and neighborhood fixed effects, restricting to buildings built after 1930, the same censoring rule as the hedonic regression. Following Diamond and Diamond (2024), we estimate the regression at the median rather than by least squares. Predicted ranks pass through the nonlinear quantile function of listed rents, so a mean-unbiased rank prediction would not deliver a mean-unbiased rent, whereas a median-unbiased one survives the monotone transformation. We use imputed rents for all properties, including those with listings, so the dependent variable in (15) is constructed identically across the sample.

The procedure relies on a single-index assumption: unobserved characteristics affect rents and assessed values only through the scalar ϵ^q in (13), and assessed values are strictly monotone in it conditional on $\mathcal{C}$, so the assessed-value rank proxies for unobserved quality. Assessed values can deviate from market prices (Berry, 2021), but the imputation uses only their ranks, and the ranks agree closely with those of sales. For properties that sold in the same year, regressing the within-year rank of assessed values on that of sale prices yields a slope of 0.83 with an R^2 of 0.69, a closer fit than the analogous log-log regression in levels, with a slope of 1.01 and an R^2 of 0.52. Unlike Diamond and Diamond (2024), we do not require the rank mapping to be stable over time, since ranks are computed within year.

B Model Appendix

B.1 Comparative Statics

This section collects the environments, formal statements, and proofs of the theoretical results in Section 4.4. We begin with the environment for Proposition 1.

Assumption 1 *We make the following assumptions for Proposition 1:*

1. *The economy lasts one period, households have Cobb–Douglas preferences ($\bar{q} = 0$) with interior quality choices, moving costs are zero, and neighborhoods have the same, exogenous amenities.*

2. *There is no renovation option.*

3. *The blueprint distribution $G(\hat{q})$ is degenerate at a single quality $\hat{q}$ common to both neighborhoods, and redevelopment builds at least one unit, $h' \in \{1, \ldots, H\}$.*

4. *The construction cost schedule $C(\hat{q}, h)$ is the same across neighborhoods.*

5. *Rents are iso-elastic in quality: $P(q, x) = \kappa(x) q^{\nu(x)}$.*

The formal counterpart of the proposition in the main text is the following.

Proposition 1 (formal statement). *Under Assumption 1, compare two neighborhoods with rent–quality elasticities $\nu_2 > \nu_1 > 0$.*

1. *Income sorting. For any two incomes $z_2 > z_1$,*

$$\frac{U(x_1, z_2)}{U(x_1, z_1)} > \frac{U(x_2, z_2)}{U(x_2, z_1)}. \tag{B.1}$$

2. *Redevelopment. Write $\Delta(x) \equiv P_2(x) - P_1(x)$ for the rent premium of neighborhood 2 at quality x, and let q_0 be the unique quality at which the two schedules cross, $\Delta(q_0) = 0$. Of two parcels with the same quality q and the same number of units h, one in each neighborhood,*

$$\log\left[\frac{p_2/(1-p_2)}{p_1/(1-p_1)}\right] = \sigma_c \left[\bar{h}\, \Delta(\hat{q}) - h\, \Delta(q)\right], \tag{B.2}$$

where p_j is the redevelopment probability of the parcel in neighborhood j and $\bar{h} \in [1, H]$ is an expected number of newly built units. Consequently $p_2 > p_1$ for every admissible blueprint quality above a finite threshold $\hat{q}^*(q, h)$. The conclusion is nonvacuous whenever the blueprint support extends above this threshold.

We next prove this proposition.

Proof of Proposition 1. We first derive how rent elasticity governs the income elasticity of indirect utility, then compare redevelopment odds across two parcels of the same quality and unit count. By Assumption 1, all quality choices below are interior.

Income sorting. The household first-order condition (4), evaluated at $P_j(q) = \kappa_j q^{\nu_j}$, gives:

$$P_j(q_j^*(z)) = \frac{\alpha z}{\alpha + (1-\alpha)\nu_j}, \qquad U(x_j, z) = C_j z^{1-\alpha+\alpha/\nu_j},$$

where $C_j > 0$ is independent of income. The income exponent decreases strictly in ν_j, proving (B.1). With zero moving costs and common amenities, neighborhood-choice probabilities satisfy $\pi_1(z)/\pi_2(z) = [U(x_1, z)/U(x_2, z)]^{\sigma_x}$, which increases strictly in z. Resident densities therefore have an increasing ratio $g_1(z)/g_2(z)$: after normalization, this ratio crosses one from below, so neighborhood 1 has a higher mean income. A flatter rent schedule makes additional quality more affordable as income rises.

Redevelopment. Write $A_j = P_j(\hat{q})$ and define the common inclusive value:

$$V(A) = \frac{1}{\sigma_c} \log \sum_{h'=1}^{H} \exp\bigl(\sigma_c[Ah' - C(\hat{q}, h')]\bigr).$$

The extreme-value shocks in (8) imply $\log[p_j/(1-p_j)] = \sigma_c[V(A_j) - hP_j(q)]$. Differentiating the inclusive value gives $V'(A) = \mathbb{E}[h' \mid A] \in [1, H]$, with the conditional probabilities in (11). Thus the mean value theorem gives $V(A_2) - V(A_1) = \bar{h}\Delta(\hat{q})$ for some $\bar{h} \in [1, H]$, proving (B.2); when $A_1 = A_2$, any such $\bar{h}$ gives the same identity.

Since $\nu_2 > \nu_1$, the rent premium $\Delta(\hat{q}) = \kappa_2 \hat{q}^{\nu_2} - \kappa_1 \hat{q}^{\nu_1}$ tends to infinity. Hence, above a finite threshold,

$$\bar{h}\Delta(\hat{q}) \geq \Delta(\hat{q}) > \max\{0, h\Delta(q)\},$$

so $p_2 > p_1$. The rent advantage of rebuilding eventually exceeds the rent forgone on the existing structure. In particular, if $q \leq q_0 < \hat{q}$, neighborhood 2 gives up weakly less rent and earns strictly more by rebuilding, so the crossing quality q_0 is already a sufficient threshold. ■

We now turn to Proposition 2, starting from the assumptions.

Assumption 2 *We make the following assumptions for Proposition 2:*

1. *Preferences are Cobb–Douglas with $\bar{q} = 0$, moving costs are zero, landlords have no renovation option, and amenities are exogenous.*

2. *Neighborhoods 0 and 1 are identical at baseline in units, quality distribution, and amenities, so that their baseline rents, resident-income distributions, and choice probabilities*

coincide, and small enough that their rents do not feed back on the rest of the city. The policy changes the housing stock of neighborhood 1 only.

3. *Quality and income have common compact supports with continuously differentiable densities bounded away from zero, $q_{\min} > 0$, and quality choices are interior except possibly at the support endpoints, where the first-order condition holds by continuity.*

4. *For each fixed $\sigma_x > 0$, equilibrium rents are locally unique and rent schedules and resident CDFs are C^1 in the policy parameters. As $\sigma_x \downarrow 0$, baseline rent schedules and resident-income densities converge in C^1 to finite limits, with the limiting density and baseline numeraire consumption uniformly bounded away from zero.*

5. *Policy changes are small, and the quality shift $\mathcal{Q}_{1,\varepsilon}$ is a C^1 family with $\mathcal{Q}_{1,0} = \mathcal{Q}_0$ and $\partial_\varepsilon \mathcal{Q}_{1,\varepsilon}(q)|_{\varepsilon=0} > 0$ at every interior quality.*

The formal statement of the proposition is as follows.

Proposition 2 (formal statement). *Under Assumption 2, there is $\bar{\sigma} > 0$ such that, for each fixed $0 < \sigma_x < \bar{\sigma}$, the following hold for sufficiently small policy changes:*

1. *Quality shift ($\mathcal{Q}_1(q) \geq \mathcal{Q}_0(q)$ for all q, strictly in the interior; $H_1 = H_0$). There is a quality q^* such that $P_1(q) \lesseqgtr P_0(q)$ as $q \lesseqgtr q^*$. Neighborhood 1 also becomes poorer: $\mathcal{Z}_1(z) \geq \mathcal{Z}_0(z)$ for all z.*

2. *Quantity expansion ($H_1 = (1 + m)H_0$ with $m > 0$; $\mathcal{Q}_1 = \mathcal{Q}_0$). Rents fall at every quality: $P_1(q) < P_0(q)$ for all q.*

We now prove this proposition.

Proof of Proposition 2. We proceed in three steps: derive the rent identity implied by assignment, establish the effects of a quality shift, and show that a quantity expansion lowers rents. For brevity, write $\gamma = \alpha/(1 - \alpha)$.

Step 1: Rent decomposition. Assortative matching assigns quality q to income $z_j(q) = \mathcal{Z}_j^{-1}(\mathcal{Q}_j(q))$. Multiplying the household first-order condition (4) by q^γ gives $(q^\gamma P_j(q))' = \gamma q^{\gamma-1} z_j(q)$. Integrating yields:

$$P_j(q) = \rho(q)P_j(q_{\min}) + q^{-\gamma} \int_{q_{\min}}^{q} \gamma s^{\gamma-1} z_j(s)\, ds, \qquad \rho(q) = (q_{\min}/q)^\gamma.$$

Thus rents reflect a common level, transmitted through $\rho(q)$, and the incomes of households assigned along the quality ladder.

Step 2: Quality shift and sorting. A dot denotes a derivative with respect to the quality-shift parameter ε at zero. Let $\mathcal{Q}$, $\mathcal{Z}$, g, and P^* denote the baseline quality CDF, resident-income CDF and density, and rent schedule. For percentile $t \in [0,1]$, write $z_t = \mathcal{Z}^{-1}(t)$ and $Q(t) = \mathcal{Q}^{-1}(t)$. At fixed quality, the direct assignment effect is:

$$d(q) = \frac{\dot{\mathcal{Q}}(q)}{g(z(q))} > 0, \qquad z(q) = \mathcal{Z}^{-1}(\mathcal{Q}(q)).$$

The inequality holds in the interior. Holding residents and the bottom rent fixed, the loss of high-quality units pushes households down the quality ladder and raises rents above the bottom quality. This is the trickle-down mechanism in reverse (Nikolakoudis, 2024; Nathanson, 2025).

Let $\dot{P}(q) = \partial_\varepsilon[P_1(q;\varepsilon) - P_0(q;\varepsilon)]|_0$ and $a(z_t) = (1-\alpha)/[z_t - P^*(Q(t))] > 0$. The envelope theorem and (7) imply:

$$\dot{\pi}_1(z_t) - \dot{\pi}_0(z_t) = -\sigma_x \pi(z_t) a(z_t) \dot{P}(Q(t)),$$

where π is the common baseline choice probability. Since unit counts are unchanged and $g(z) = \ell(z)\pi(z)/H$, with ℓ the citywide income density and H baseline units, integration gives:

$$\dot{\mathcal{Z}}_1(z_t) - \dot{\mathcal{Z}}_0(z_t) = -\sigma_x B(t), \qquad B(t) = \int_0^t a(z_s)\dot{P}(Q(s))\, ds, \qquad B(0) = B(1) = 0.$$

The last condition is neighborhood clearing. Differentiating assignment gives the total income response $d(Q(t)) + \sigma_x B(t)/g(z_t)$. Normalize the rent response as $W(t) = \dot{P}(Q(t))/\rho(Q(t))$. Combining assignment with Step 1 yields:

$$B' = wW, \qquad W' = b\left[d(Q(t)) + \frac{\sigma_x B}{g(z_t)}\right],$$

where $w(t) = a(z_t)\rho(Q(t)) > 0$ and $b(t) = \gamma q_{\min}^{-\gamma} Q(t)^{\gamma-1} Q'(t) > 0$. The first equation measures the migration response to rents; the second measures how the resulting income composition feeds back into the rent gradient.

We first show that $B < 0$ in the interior. Eliminating W gives:

$$(B'/w)' - \frac{\sigma_x b}{g(z_t)} B = b\, d(Q(t)) > 0, \qquad B(0) = B(1) = 0.$$

If B were nonnegative anywhere in the interior, it would attain a nonnegative interior maximum. There $B' = 0$ and $B'' \leq 0$, making the left side nonpositive, a contradiction. Thus the resident CDF rises to first order: mobility makes the treated neighborhood poorer and offsets part of the direct increase in assigned income.

To establish the rent crossing, first omit the composition term from the equation for W'. Define its cumulative direct effect and its clearing-weighted mean:

$$r(t) = \int_0^t b(s) d(Q(s))\, ds, \qquad \bar{r} = \frac{\int_0^1 w(t) r(t)\, dt}{\int_0^1 w(t)\, dt}.$$

The direct effect r is strictly increasing, and clearing requires a downward level adjustment $\bar{r}$. With composition adjustment restored, integration of W' and $\int_0^1 wW\, dt = 0$ gives:

$$W = r - \bar{r} + E, \qquad E(t) = J(t) - \frac{\int_0^1 w(s) J(s)\, ds}{\int_0^1 w(s)\, ds}, \qquad J(t) = \sigma_x \int_0^t \frac{b(s) B(s)}{g(z_s)}\, ds.$$

The regularity conditions make $w, b, 1/g$ uniformly bounded, with $\int w$ bounded away from zero. Hence $B' = wW$ and the formula for J imply $\|E\|_{C^1} \leq C\sigma_x \|W\|_\infty$. Since r is uniformly bounded, substituting the expression for W and absorbing the last term for sufficiently small σ_x gives $\|E\|_{C^1} = O(\sigma_x)$.

The limiting function $r - \bar{r}$ is negative at zero, positive at one, and strictly increasing in the interior. Its unique zero is interior and has positive derivative. The C^1 bound therefore preserves its signs away from that zero and its positive derivative nearby. Thus W, and hence $\dot{P}$, crosses zero exactly once, from below. For each fixed sufficiently small σ_x, the C^1 policy expansion transfers this crossing to $P_1 - P_0$ for sufficiently small $\varepsilon > 0$.

It remains to make the sorting comparison uniform over incomes. Let $F_\varepsilon(t) = \mathcal{Z}_1(z_t; \varepsilon) - \mathcal{Z}_0(z_t; \varepsilon)$. Its derivative at zero is $\dot{F}(t) = -\sigma_x B(t) > 0$ in the interior. Moreover, $\dot{F}'(0) = -\sigma_x w(0) W(0) > 0$ and $\dot{F}'(1) = -\sigma_x w(1) W(1) < 0$. The C^1 expansion $F_\varepsilon / \varepsilon \to \dot{F}$ preserves these derivative signs near the endpoints and positivity on the remaining compact interior. Since $F_\varepsilon(0) = F_\varepsilon(1) = 0$, we obtain $F_\varepsilon(t) > 0$ throughout the interior. This proves part 1: cheaper low-quality housing attracts poorer residents, while the loss of high-quality units raises rents at the top.

Step 3: Quantity expansion. Hold $\mathcal{Q}_1 = \mathcal{Q}_0 = \mathcal{Q}$ and let $H_1 = (1+m) H_0$. Write:

$$P_m(q) = \partial_m [P_1(q; m) - P_0(q; m)]|_0\,, \qquad Y(t) = Q(t)^\gamma P_m(Q(t)).$$

Let $D(t) = \partial_m[\mathcal{Z}_1(z_t; m) - \mathcal{Z}_0(z_t; m)]|_0$. Differentiating assignment and resident densities gives:

$$Y' = -\lambda D, \qquad D' = -1 - cY, \qquad \lambda(t) = \frac{\gamma Q(t)^{\gamma-1}Q'(t)}{g(z_t)} > 0, \quad c(t) = \sigma_x a(z_t)Q(t)^{-\gamma} > 0.$$

The term -1 comes from dividing treated residents by the expanded stock $(1+m)H_0$; the term $-cY$ is the migration response to rents. Since $D(0) = D(1) = 0$, eliminating D yields:

$$(Y'/\lambda)' - cY = 1, \qquad Y'(0) = Y'(1) = 0.$$

At a nonnegative global maximum, $Y' = 0$ and $Y'' \leq 0$, including at an endpoint by the boundary conditions. The left side would then be nonpositive, a contradiction. Thus $Y < 0$, so $P_m(q) < 0$ at every quality. Compactness and the uniform policy expansion imply $P_1(q; m) - P_0(q; m) < 0$ for sufficiently small $m > 0$. Rents must fall throughout the ladder to attract the households needed to fill the additional units. This proves part 2 and completes the proof. ■

Finally, the next proposition formalizes how the two anti-gentrification policies discussed in Section 4.4—a teardown tax, which charges a fee T for demolishing a structure, and an anti-deconversion ordinance, which prohibits redevelopment that reduces the number of units in a building—curb redevelopment. We work in the one-period setting of Assumption 1.

Proposition 3 (The effects of anti-gentrification policies on redevelopment) *Fixing the rent function, the following hold, where p denotes a parcel's redevelopment probability before the policy.*

1. Teardown tax. *For any $T > 0$, the teardown tax reduces the redevelopment probability to $p(T) < p$:*
$$\log \frac{p(T)}{1 - p(T)} = \log \frac{p}{1-p} - \sigma_c T. \tag{B.3}$$

2. Anti-deconversion. *The anti-deconversion ordinance reduces the redevelopment probability to $p^{AD} \leq p$, with strict inequality whenever the ordinance binds, i.e., $S(\sigma_c) < 1$:*
$$\log \frac{p^{AD}}{1 - p^{AD}} = \log \frac{p}{1-p} + \log S(\sigma_c), \tag{B.4}$$
where $S(\sigma_c) = \sum_{h' \geq h} e^{\sigma_c \pi^R(h')} / \sum_{h' \in \mathcal{H}} e^{\sigma_c \pi^R(h')} \in (0, 1]$ is the conditional probability, given redevelopment before the policy, of choosing a non-unit-reducing option.[48]

[48]Following Kalouptsidi et al. (2021), we model the ordinance as a state-contingent removal of redevel-

The proof of the proposition follows.

Proof of Proposition 3. With shocks of dispersion σ_c on inaction and each unit choice, write the redevelopment odds as

$$O \equiv \frac{p}{1-p} = \frac{\sum_{h' \in \mathcal{H}} e^{\sigma_c \pi^R(h')}}{e^{\sigma_c \pi^N}}.$$

Teardown tax. The tax replaces each redevelopment payoff by $\pi^R(h') - T$ and leaves inaction unchanged. Thus $O(T) = e^{-\sigma_c T} O$, which gives (B.3); since $p = O/(1+O)$ increases in O, $p(T) < p$ for $T > 0$ and $\partial p / \partial T = -\sigma_c p(T)[1 - p(T)]$.

Anti-deconversion. The ordinance retains options with conditional probability share S, so $O^{AD} = SO$ and (B.4) follows. If $S < 1$, then $p^{AD} < p$; if $S = 1$, the ordinance is nonbinding and $p^{AD} = p$. ■

B.2 Solution Algorithm

We discretize the quality space $\mathbb{Q}$ into a grid $\{q_1, q_2, \ldots, q_{N_q}\}$ with $q_1 = \bar{q}$ and levels evenly spaced in logs, as described in Section 5.2.1. Quality depreciates stochastically: a structure falls one grid level with probability π_δ and otherwise retains its quality, with π_δ chosen so that the expected decline in log quality per period matches δ. Under the assumption of perfect foresight, the model is solved iteratively in the following steps:

1. Start by solving for the long run steady state equilibria. The policies we report in this paper are temporary, so this is the baseline steady state.

2. Solve for the dynamic equilibrium up to a pre-specified terminal period T. Assume all value functions at the terminal period are given by the steady state value functions solved above. To do this, proceed in the following way:

   (a) Guess the entire stream of the rent function $\{P_t(q, x)\}_{\forall t, q, x}$.

   (b) Given the rent function, solve the household's problem by equations (2) and (5). This generates the housing demand $\{L_t(q, x, z)\}_{\forall t, q, x, z}$.

   (c) Given the rent function, solve the landlord's problem in reverse order (from $t = T$ to $t = 1$) in the following steps:

opment options: conditional on the current unit count h, choices with $h' < h$ are infeasible. Their result identifies the choice response when the policy only removes options and leaves the payoffs and state transitions of feasible options unchanged; welfare and general-equilibrium effects require additional model structure.

i. In period T, set the $T+1$ continuation $V_{T+1}(s_{iT}, \hat{q}_{iT})$ to be the steady state value function solved in Step 1.

ii. In period $t \le T$, solve the problem given by equation (8). The value functions $V_{it+1}(\cdot, \cdot)$ are solved from the landlord's problem in the period $t+1$.

A. Calculate the value of inaction, net of the idiosyncratic shock:

$$V_{it}^{N}(s_{it}) = P_t(q_{it}, x)h_{it} + \beta \mathbb{E}\, V_{i,t+1}\left(D(q_{it}), h_{it}, \hat{q}_{i,t+1}, \vec{\xi}_{i,t+1}\right),$$

where the expectation is taken over the depreciation draw, the next blueprint draw, and the next period's shocks.

B. Calculate the value of renovation, which raises quality while preserving the unit count:

$$\begin{aligned} V_{it}^{RN}(s_{it}) = &- C^{RN}(q_{it}, h_{it}) + P_t\big((1+\varrho)q_{it}, x\big)h_{it} \\ &+ \beta \mathbb{E}\, V_{i,t+1}\left(D\big((1+\varrho)q_{it}\big), h_{it}, \hat{q}_{i,t+1}, \vec{\xi}_{i,t+1}\right). \end{aligned}$$

C. Calculate the value of redeveloping at the blueprint quality $\hat{q}_{it}$ with h units,

$$V_{it}^{R}(\hat{q}_{it}, h) = -C_x(\hat{q}_{it}, h) + P_t(\hat{q}_{it}, x)h + \beta \mathbb{E}\, V_{i,t+1}\left(D(\hat{q}_{it}), h, \hat{q}_{i,t+1}, \vec{\xi}_{i,t+1}\right),$$

and aggregate over the unit choice, $V_{it}^{R}(s_{it}, \hat{q}_{it}) = \frac{1}{\sigma_c} \log \sum_{h \in \mathcal{H}(h_{it})} \exp\left(\sigma_c V_{it}^{R}(\hat{q}_{it}, h)\right)$. The feasible set $\mathcal{H}(h_{it})$ is all of $\mathcal{H}$ absent the anti-deconversion ordinance and $\{h : h \ge h_{it}\}$ under it. Because redevelopment replaces the existing structure, the payoff depends on the current state only through the added cost φ_x of demolishing a multifamily structure and, under the ordinance, through the feasible set.

D. Calculate the probabilities of each action:

$$P^{j}(s_{it}, \hat{q}_{it}) = \frac{\exp\left(\sigma_c V_{it}^{j}(s_{it}, \hat{q}_{it})\right)}{\sum_{j' \in \{N, RN, R\}} \exp\left(\sigma_c V_{it}^{j'}(s_{it}, \hat{q}_{it})\right)}, \quad j \in \{N, RN, R\}.$$

(d) Given the action probabilities $\{P^N, P^{RN}, P^R\}$ and unit choices solved in Step 2(c), and the initial distribution of parcels over states $\{s_{i0}\}_i$, iterate the distribution forward from $t = 1$ to $t = T$ and aggregate to housing units by quality, $H_t(q, x)$. Parcels that take no action carry their units to the depreciated quality; parcels that renovate move up one grid level with probability π_{RN} and otherwise stay;

parcels that redevelop move to the blueprint quality with their chosen unit count.

(e) Check the market clearing condition:

$$\sum_{z \in \boldsymbol{Z}} L_t(q,x,z) = H_t(q,x), \quad \forall t \in \{1,2,...,T\}, q \in \boldsymbol{Q}, x \in \boldsymbol{X}$$

i. if the market clearing condition is satisfied, stop the iteration.

ii. if not satisfied, update the rent function $\{P_t(q,x)\}_{\forall t,q,x}$ in a way proportional to excess demand; and go back to Step 2(a).

C Model Estimation Details

In this appendix, we provide supplementary details for the empirical implementation of our model.

C.1 Quality and Density in New and Old Buildings

We report two important sets of targeted moments that our model matches, inspired by the empirical evidence in Section 3.

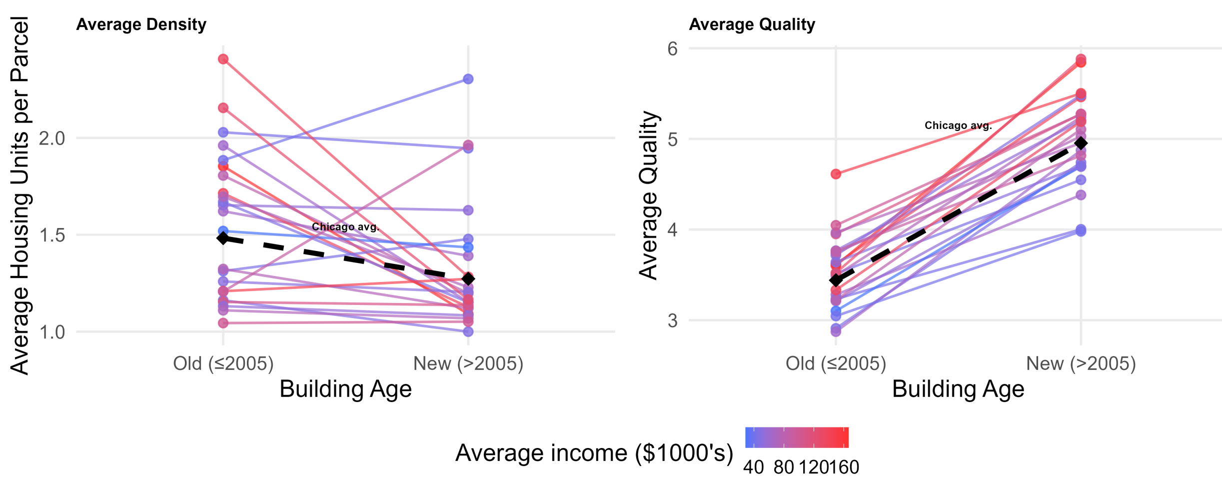


Figure C.1: Density and Quality in New and Old Structures by Neighborhood Group

Notes: The figure plots differences in average housing units per building and in average quality between old and new structures, by neighborhood group. New buildings are those built after 2005. Neighborhood groups are shaded by average income and connected across building-age bins by solid lines; Chicago averages are reported in black. High-income neighborhoods are associated with lower average density and greater quality upgrading in new structures.

C.2 Details on Calibrating the Redevelopment Elasticity

We calibrate σ_c by reproducing the estimated demolition response to the ordinances in the model, treating the 606 Trail and Pilsen areas as two small neighborhoods. In the data, the treated parcels are the assessed residential buildings whose centroids fall inside the ordinance boundaries, taken from the most recent assessment of each parcel between 2016 and 2020. Their model counterpart is a distribution of parcels over quality and unit counts, $\{m_s\}_{s=(q,h)}$, which we build from the same records. We place each area's housing units on the model's quality grid by quantile: a unit is assigned the grid point whose percentile in the calibrated citywide quality distribution equals the unit's percentile in the pooled estimated quality distribution of the two areas, and a small mass is placed at the minimum quality to represent vacant or derelict parcels. We combine this quality distribution with the area's distribution of unit counts among existing buildings, and scale the result so that the model's total number of housing units equals the area's. This initial distribution is then adjusted during calibration so that the model's overall quality distribution matches the area's.

We solve the landlord's problem on these parcels at a three-year period so that the redevelopment decision spans the policy window, and we measure the moments that discipline the supply side in the main calibration of Section 5.2.3 over the three years before the policy: the redevelopment rates of single-family and multifamily structures, the renovation rate, the distribution of parcels across unit counts, and the quality and density of buildings built after 2005 relative to older ones. We convert the annual discount and depreciation rates to their three-year equivalents and discount flow rents within the period as a three-year annuity. We assume the rent function is unchanged by the policy, consistent with the null effect on housing prices in the boundary design (Figure D.8). We also assume that landlords expect the ordinances to expire after three years, as announced, and that the policy does not affect the future value of a parcel once it's lifted.

Because construction costs depend on the period length, we recalibrate them at the three-year horizon for every candidate σ_c. Given σ_c, we choose the fixed cost schedule over blueprint qualities, the variable cost schedule over unit counts, the renovation cost, the multifamily demolition premium, and the distribution of treated parcels over quality and unit counts to match the three-year moments above. We then hold these costs and rents fixed, impose the surcharge on the fixed cost of demolition and remove unit-reducing projects from the choice set, and compute each parcel's redevelopment probability with and without the policy.

We match the Poisson quasi-maximum-likelihood version of the model's policy effect to its empirical counterpart. In the data, demolition is rare for every parcel, so the Poisson estimate of the policy effect is very close to the logit estimate reported in Section 5; the choice of estimator matters only for the model moment. Let m_s denote the mass of treated parcels in state $s = (q, h)$, and let p_s^0 and p_s^1 denote the redevelopment probabilities without and with the policy, integrated over blueprint draws. The model moment is

$$\beta^P = \log\left(\frac{\sum_s m_s\, p_s^1}{\sum_s m_s\, p_s^0}\right), \tag{C.1}$$

with the sums pooled across the two areas. This is the log proportional change in expected demolitions, the estimand of a Poisson quasi-maximum-likelihood regression of demolition on treatment (Wooldridge, 2023). We choose σ_c so that β^P equals the empirical coefficient. The logit estimand instead averages parcel-level changes in log odds. To first order, $\beta^L \approx \sum_s w_s \left[\log \frac{p_s^1}{1-p_s^1} - \log \frac{p_s^0}{1-p_s^0}\right] / \sum_s w_s$, with variance weights $w_s = m_s p_s^0(1 - p_s^0)$ evaluated at the baseline probabilities.

We target β^P because it is far less sensitive than β^L to heterogeneity in the policy's effect across parcels. Because the anti-deconversion ordinance only applies to multifamily structures, the combined policies' effect on the log odds of redevelopment is heterogeneous across parcels and larger for multifamily parcels, and in a model that abstracts from many options available to owners of aging buildings, this gap is wide. A multifamily parcel whose redevelopment probability decreases significantly contributes a very large negative number to β^L but only its own share of demolitions to β^P. In the data, the Poisson coefficient aggregates such heterogeneous responses into a log ratio of demolition counts, and the logit we estimate, which assumes a homogeneous treatment effect, delivers the same aggregate because the demolition rate is low for every parcel. Matching β^P therefore compares the model with the data on the object the empirical design identifies.

C.3 Details on Calibrating the Supply Side of the Model out of Steady State

A key feature of the calibration procedure is that we do *not* target the model's steady-state housing stock. Instead, we calibrate directly to the observed 2020 cross-section of Chicago parcels. Targeting the entire transition path during calibration requires solving the full spatial general equilibrium at every candidate parameter vector. This is a computationally prohibitive task given the high-dimensional state space and the need to iterate across all 23 neighborhood groups simultaneously. By decoupling the supply-side calibration from the equilibrium solver, we match the actual distribution of housing units and qualities in the data while circumventing the fixed-point computation over future equilibrium prices that targeting a rational-expectations equilibrium would entail.

To evaluate landlord value functions off the steady state, we require an assumption about price expectations. We adopt *steady-state price expectations*: landlords discount future payoffs using the current hedonic rent schedule $P(q,x)$, treating it as if it will persist indefinitely. Formally, the value of holding a unit of quality q in neighborhood x is computed under the assumption that rents do not change over the planning horizon. This is a useful approximation in dynamic discrete-choice models of durable-good investment when the economy changes slowly, and it substantially reduces the computational burden of the calibration by allowing value functions to be evaluated without solving for future equilibrium prices. Importantly, we do *not* impose steady state expectations when solving for any equilibria; we only use them to calibrate the model.

The assumption of steady-state price expectations is well-grounded in the economics of this setting. The housing stock evolves slowly: annual quality depreciation is approximately $\delta = 0.34\%$ per year, implying that the distribution of housing quality drifts only modestly within a single model period of ten years. The counterfactual analysis in Section 6 corroborates this: even under large policy shocks, the model economy takes multiple decades to converge to a new steady state. Landlords forming expectations over the planning horizon therefore face a price environment that changes little relative to the investment decisions at hand, validating the steady-state expectations approximation as a source of negligible distortion to the true calibrated parameters. We confirm that the first-period rents by quality segment and incomes match calibrated counterparts closely, with maximum deviations of a few percentage points.

C.4 Construction Costs

We use each of the estimated and internally calibrated parameters of the housing production function to construct model-based measures of construction costs. We report these estimates for each neighborhood and blueprint quality level in Figure C.2.

C.5 Model Fit on Untargeted Moments

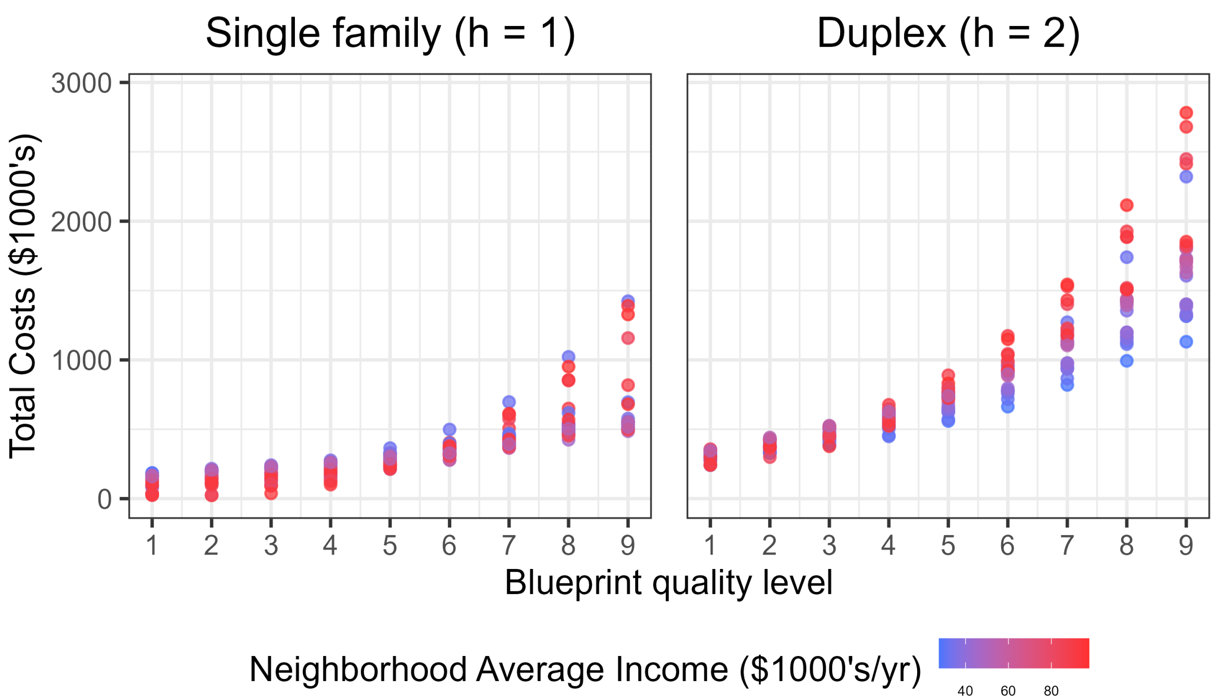


Figure C.2: Model-Based Construction Costs by Neighborhood and Blueprint Quality

Notes: The figure reports model-based construction costs, constructed from the estimated and internally calibrated parameters of the housing production function in Section 5. The support of the blueprint distribution is the quality grid described in Section 5.2.1, excluding the minimum quality level; costs are plotted for representative low, middle, and high quality levels of that grid. Construction costs for the highest-quality blueprints are exponentially greater in all neighborhoods, and higher-income neighborhoods tend to face greater costs for high-quality blueprints. Costs for duplexes tend to be higher, especially in higher-income neighborhoods.

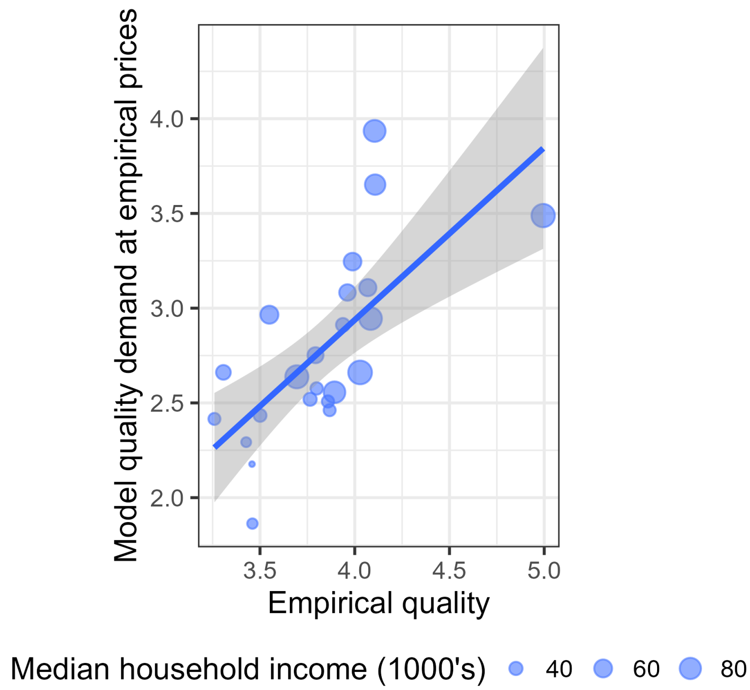


Figure C.3: Model Fit on Untargeted Moments: Average Housing Quality by Neighborhood Group

Notes: The figure plots empirical quality, measured from the hedonic regression in Section 5, against the average quality demanded in the model at empirical prices, by neighborhood group. Marker size is proportional to median household income. The R^2 of this regression is 50%, indicating that the model captures meaningful variation in equilibrium housing quality across neighborhood groups. Model quality levels are lower than empirical quality levels on average because we target rent spending shares from the ACS, which consistently reports lower rents than RentHub listings. Because quality is an index with no cardinal interpretation, this level difference has no consequence for model outcomes.

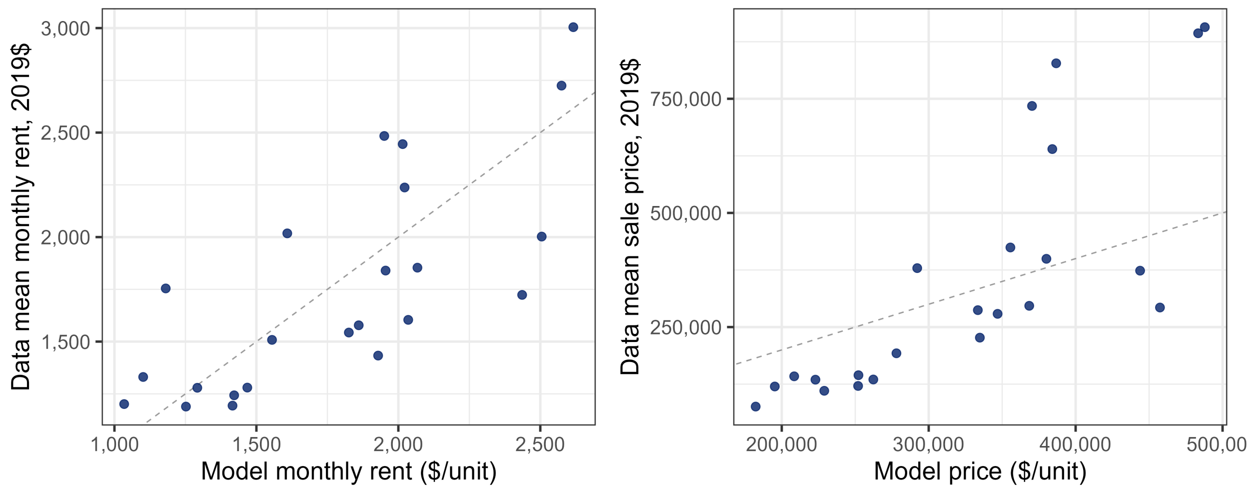


Figure C.4: Plots of equilibrium model average rent and average housing prices against empirical counterparts measured from Renthub and transactions data. Variation is across neighborhoods. R^2 for rent and prices are 60% and 65%, respectively.

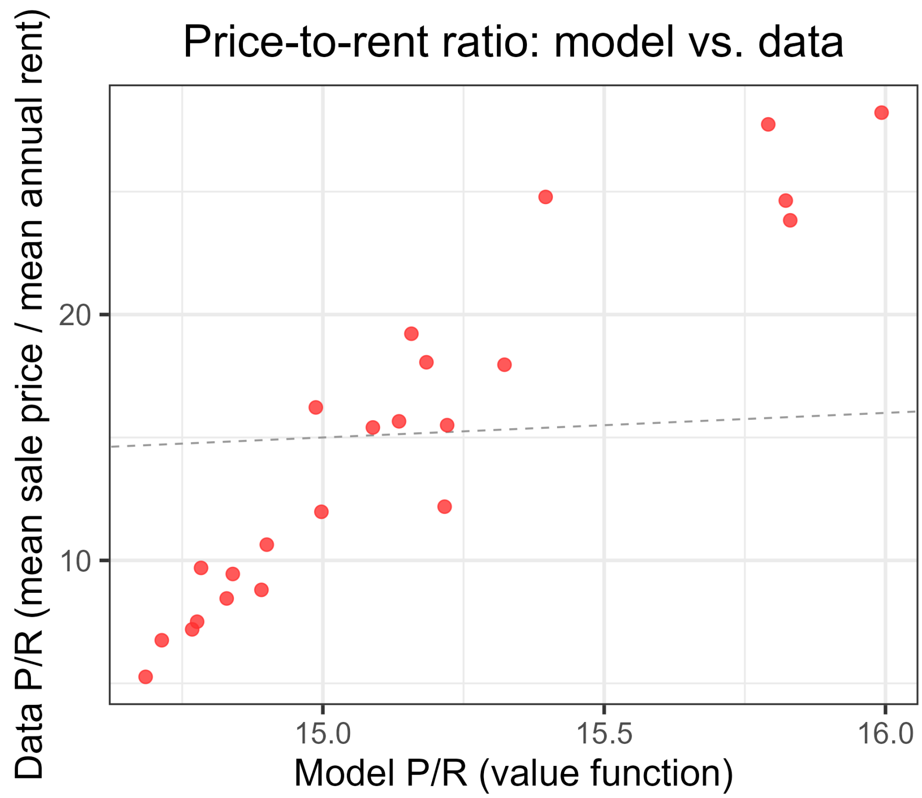


Figure C.5: Plots of equilibrium model price-to-rent ratio against data price-to-rent ratios. Variation is across neighborhoods. While our model captures neighborhood variation in price-to-rent (PTR) with high correlation, across neighborhood variation in PTR is small in the model. This is because we do not capture potential variation in discount factors across space that could be rationalized by variation in rental yield risk.

D Policy Appendix

This appendix documents the two 606–Pilsen ordinances whose effect on redevelopment identifies σ_c in Section 5.2.3. We discuss the institutional setting and the details of our empirical designs.

D.1 Policy Context

The City Council adopted the two pilot ordinances, in force from April 2021 to April 2024, in response to a decade in which multifamily buildings in the 606 Trail and Pilsen were torn down and replaced by higher-end single-family homes. Figure D.1 maps the two areas, which together hold roughly 4 percent of Chicago's housing stock. The 606–Pilsen Demolition Permit Surcharge sets a fee at the greater of $15,000 per permit and $5,000 per demolished residential unit, with the revenue directed to the Chicago Community Land Trust, which subsidizes affordable homeownership.[49] The Anti-Deconversion Ordinance bars any project from reducing the number of housing units on a parcel. The two ordinances aim to curb redevelopment and gentrification in the treated neighborhoods.

Both neighborhoods were suffering from high demolition rates and gentrification before the ordinances, which is why they were selected. Figure D.2 shows that demolition and construction rates were well above the city average rates through 2020, and that average rents and sale prices grew faster than in the rest of the city. After April 2021, demolition and construction fall, and rent and price growth drops below the city average. These movements alone cannot be read as the effect of the ordinances: they coincide with the pandemic and with a citywide slowdown, which is what the two designs below are built to difference out.

A demolition ban preceded the ordinances in part of the 606 area, running from February 2020 until the surcharge replaced it. It could have depressed teardowns in the treated area during the sample's last pre-policy period. The event studies below show no such effect: the pre-policy coefficients are insignificant in both designs, including for 2018–2020. Had the ban bound, we would expect demolitions to rebound once it was lifted, which would make us underestimate the policy effect.

D.2 Samples and Specification

The two designs estimate the same equation (16) and differ in where the control parcels come from: the boundary design draws them from the 500-meter ring outside the policy boundary, the matched design from block groups elsewhere in Chicago that resemble the treated ones.

[49] The surcharge is waived if the demolition addresses imminently hazardous housing conditions, or if the replacement building reserves at least half of its units for low-income residents.

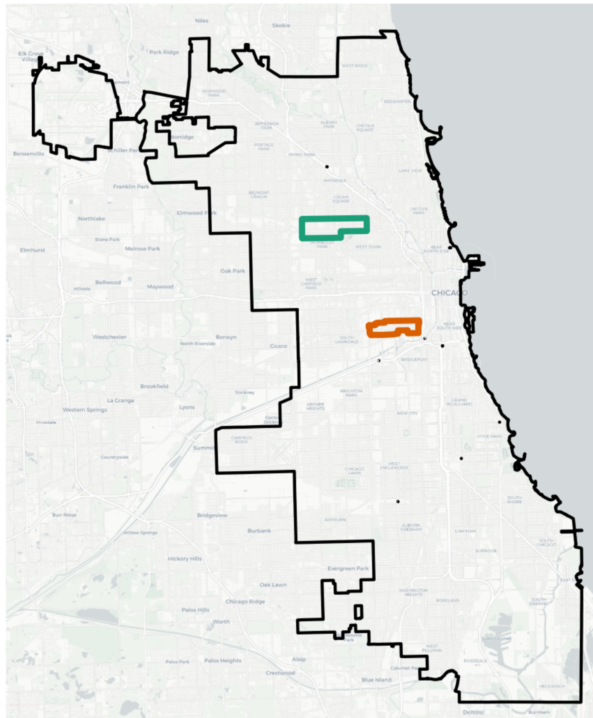
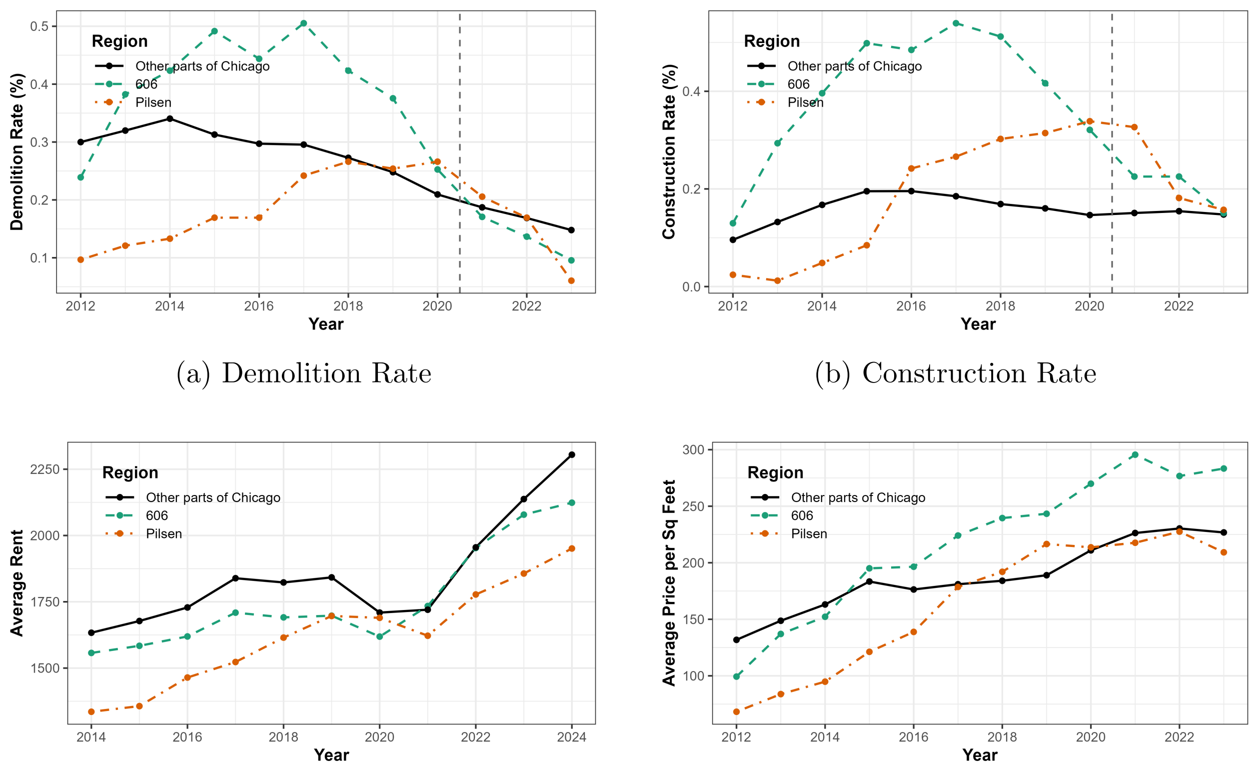


Figure D.1: Policy areas: The 606 and Pilsen neighborhoods

Note: The solid line represents the boundary of the city of Chicago. The upper-left area is 606-Trail; the other area is Pilsen.

They are complementary. The boundary controls share the treated area's housing stock and demand conditions but sit close enough to be affected by spatial spillovers of the policy; the matched controls are free of that contamination but may face different local shocks even conditional on the geographic controls.

The unit of observation is the 2005 parcel from the assessment file used throughout the paper, and the outcome is an indicator for a demolition, construction, renovation, or deconversion permit in a three-year period. We estimate at three-year periods because permits are rare at the parcel level: 0.8 percent of parcels in the boundary sample receive a demolition permit in a period (Table D.4). The rent regressions use the RentHub listings described in Section 2. We geocode each listing and match it to an assessment record by address, taking building characteristics from the assessment record when the two conflict, drop listings with rents or floorspace below the 1st or above the 99th percentile, and collapse the remaining listings to the unit-year level by averaging posted rents within a year. The sale-price regressions use the transaction deeds from the Cook County Recorder of Deeds, restricted to arm's-length sales: we exclude sales below \$10,000, special deed types such as quit-claim or



(a) Demolition Rate

(b) Construction Rate

(c) Average Rent

(d) Average Price per Square Foot

Figure D.2: Demolition, Construction, Rent, and Price Trends in the 606 and Pilsen Neighborhoods

executor deeds, and duplicate records, and match each transaction to its assessment record by PIN. We estimate permit outcomes by logit and the rest by ordinary least squares, throughout with Conley spatial standard errors with a 100-meter cutoff and 2018–2020 as the reference period.

We control for parcel fixed effects in the permit regressions. The rent regressions use the unit identifier from the assessment records and control for unit fixed effects. The sale-price regressions control for bedrooms, bathrooms, building age, unit count, and floorspace instead, since repeated transactions of the same property within this window are rare and selected. In the boundary design we include the area-by-period fixed effects α_{xt}, which let the 606 Trail and Pilsen follow separate trajectories, and the period-specific linear controls in longitude and latitude $F_t(LON_i, LAT_i)$.

Table D.1 compares the housing stock within 500 meters on the two sides of the policy boundary, using the assessment, transaction, and rental samples. The two sides are well balanced: no assessment characteristic and neither sale-price measure differs significantly across the boundary. The exceptions are that transacted buildings on the treated side have fewer units, and that rental listings there are somewhat smaller, older, and cheaper, with

Table D.1: Balance Test between the Buffer Areas within and outside the Treatment Boundary

	Treated Area		Control Area		Difference	
Variable	Mean	SD	Mean	SD	Estimate	SE
Panel A: Assessment data (2020)						
Bedrooms	4.57	(2.35)	4.38	(2.19)	0.20	(0.24)
Bathrooms	2.51	(1.26)	2.41	(1.17)	0.10	(0.15)
Unit Sq. ft.	2423.36	(1433.72)	2337.09	(1369.41)	86.28	(207.49)
Land Sq. ft.	3201.29	(899.91)	3198.47	(1104.07)	2.81	(233.91)
Single Family	0.34	(0.47)	0.41	(0.49)	-0.07	(0.09)
Building Units	2.14	(1.20)	1.97	(1.16)	0.16	(0.23)
Build Year	1909.87	(39.14)	1916.59	(38.35)	-6.72	(8.35)
Panel B: Transaction data (2015–2020)						
Bedrooms	4.05	(1.94)	3.67	(1.83)	0.38	(0.25)
Bathrooms	2.50	(1.08)	2.39	(1.06)	0.11	(0.09)
Unit Sq. ft.	2120.08	(953.13)	1982.80	(938.16)	137.28	(142.77)
Land Sq. ft.	3498.74	(1256.97)	3702.12	(1594.08)	-203.38	(256.27)
Single Family	0.43	(0.49)	0.48	(0.50)	-0.05	(0.11)
Building Units	2.66	(2.71)	3.35	(3.53)	-0.69***	(0.14)
Build Year	1927.54	(51.60)	1938.22	(50.49)	-10.68	(10.39)
log(Sale Price)	12.79	(0.62)	12.84	(0.65)	-0.05	(0.03)
log(Sale Price / Unit Sq. ft.)	5.23	(0.58)	5.35	(0.59)	-0.12	(0.09)
Panel C: Rental data (2015–2020)						
Bedrooms	2.18	(0.87)	2.15	(0.75)	0.03	(0.06)
Bathrooms	1.28	(0.51)	1.40	(0.54)	-0.11**	(0.05)
Unit Sq. ft.	1073.35	(353.25)	1120.49	(329.16)	-47.14*	(26.83)
Build Year	1897.47	(24.89)	1907.40	(29.26)	-9.93*	(5.49)
log(Rent)	7.35	(0.31)	7.47	(0.33)	-0.13***	(0.04)
log(Rent / Unit Sq. ft.)	0.42	(0.29)	0.49	(0.28)	-0.07***	(0.02)

Note: Standard errors are clustered by treatment-area group. * p<0.1, ** p<0.05, *** p<0.01.

rents about 13 log points lower.

As discussed in Baum-Snow and Ferreira (2015), the policy may generate spatial spillovers to the area outside the boundary, which would cause the boundary difference-in-differences to overestimate the policy effect. Two pieces of evidence indicate that such spillovers are limited. First, demolition and construction rates just outside the boundary do not rise after the ordinances, and the ring's demolition rate falls at roughly the citywide pace (Figure D.3). Second, contractors that worked only inside the treated area rarely relocate to the ring after the policy; most move elsewhere in Chicago or stop pulling permits altogether (Table D.3).

The matched design pairs each treated block group with up to three untreated block groups at least one kilometer from the boundary by propensity-score matching. We match block groups by log population density, log median rent, log median home value, log median household income, and median building age, averaged over 2013–2020, and by their demo-

Table D.2: Balance Test: Propensity-Score Matched Block Groups

	Treated	Matched	Difference	
Variable	Mean	Mean	Diff	p-value
Log population density	8.848	9.046	-0.198	0.161
Log median rent	6.882	6.929	-0.047	0.226
Log median home value	12.457	12.547	-0.089	0.158
Log median household income	10.768	10.842	-0.074	0.288
Median building age	73.3	73.5	-0.2	0.878
Demolition permit rate	0.004	0.004	0.000	0.743
Construction permit rate	0.005	0.003	0.001	0.153
Renovation permit rate	0.023	0.024	-0.002	0.659

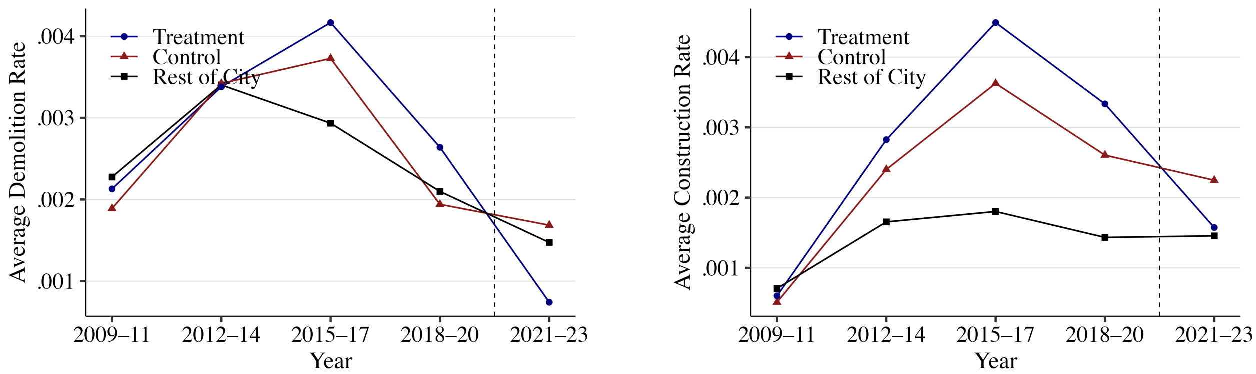
Note: Means for the 42 treated block groups and their 122 propensity-score matched controls, with p-values from two-sample t-tests. Controls are chosen by nearest-neighbor matching without replacement on a logit propensity score, with a caliper of 0.2 and up to three controls per treated block group; 40 treated block groups find three controls and two find one. Characteristics are averaged over 2013–2020, building age over 2013–2017. Permit rates are the share of a block group's 2005 parcels receiving a permit of the given type.

lition, construction, and renovation permit rates in each of the four pre-policy periods. All 42 treated block groups find controls, for 122 control block groups in total, and none of the eight mean differences between the two groups reaches the 10 percent level (Table D.2). We estimate (16) on this sample with the same parcel fixed effects and coordinate controls as in the boundary design, but with period fixed effects in place of α_{xt}: we do not control for the area-by-period fixed effects because the matched controls lie outside the two areas that x indexes.

Each design assumes that its control parcels would have followed the treated parcels' path absent the ordinances, and each could fail for a different reason. Pre-policy demolition coefficients are insignificant in both designs, as are the renovation and deconversion coefficients; the exception is construction in the matched design, which shows an upward trajectory of the treated parcels until the policy.

D.3 Effects on Building Permits

Demolitions fall in both designs (Figures D.4 and D.5). The post-policy logit coefficient is -1.23 in the boundary design and -0.73 in the matched design, both significant at the 5 percent level or better, which correspond to reductions in the demolition rate of about 71 and 52 percent. The two are not statistically distinguishable, and Section 5.2.3 targets a coefficient of -1, between them. Construction permits fall by about as much, with coefficients of -0.93 in the boundary design, significant at the 5 percent level, and -0.43 in the matched design,



(a) Average Demolition Rate

(b) Average Construction Rate

Figure D.3: Average Demolition and Construction Rates within 500-meter Buffer Zones

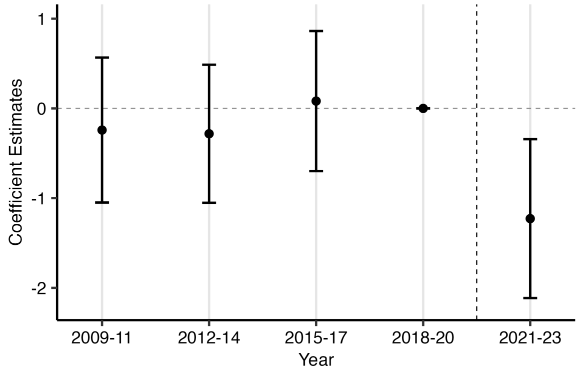
Note: The figure shows the average demolition rate (Panel (a)) and construction rate (Panel (b)) within the 500-meter buffer zones inside and outside the policy boundary, compared with the citywide average. Rates are measured as the share of 2005 parcels issued each type of building permit in each three-year period.

Table D.3: Spatial Reallocation of Contractors

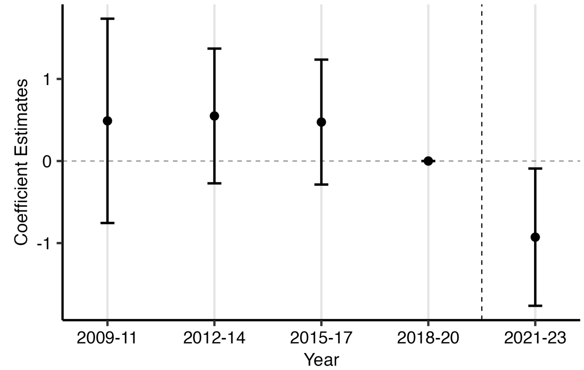
	Post-Policy Location					
Pre-Policy Location	Treated	Control	Both	Other City	Exit	Total
Panel A. Demolition–wrecking contractors						
Treated Only	0 (0%)	1 (8%)	0 (0%)	6 (50%)	5 (42%)	12
Control Only	0 (0%)	0 (0%)	0 (0%)	5 (45%)	6 (55%)	11
Both	4 (13%)	3 (10%)	2 (7%)	15 (50%)	6 (20%)	30
Total	4	4	2	26	17	53
Panel B. Construction–general contractors						
Treated Only	10 (11%)	3 (3%)	0 (0%)	16 (18%)	61 (68%)	90
Control Only	0 (0%)	3 (4%)	0 (0%)	18 (26%)	49 (70%)	70
Both	5 (11%)	3 (7%)	7 (15%)	18 (39%)	13 (28%)	46
Total	15	9	7	52	123	206

Note: The table tracks contractors operating in the 500m buffer zones around the treatment boundary before and after the demolition surcharge policy. Panel A covers demolition–wrecking contractors and Panel B construction–general contractors. New entrants, active post-policy only and absent from the buffer pre-policy, number 6 in Panel A (2 treated only, 4 control only, 0 both) and 51 in Panel B (20 treated only, 29 control only, 2 both). Column headings "Treated" and "Control" denote post-policy activity in that area only. "Other City" = entity has no post-policy permits within the 500m buffer but appears elsewhere in Chicago. "Exit" = no permits anywhere in Chicago post-policy. Cell values show count (% of row total).

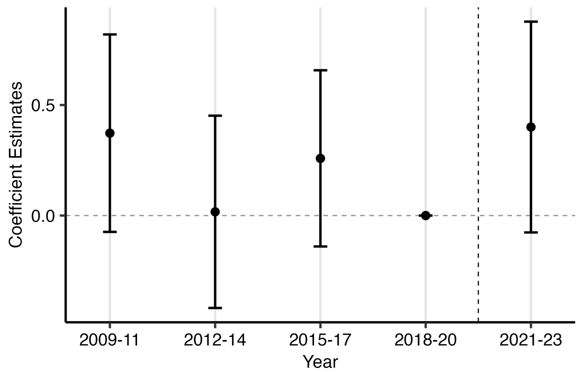
where it is not distinguishable from zero. The joint decline in demolition and construction permits reflects a reduction in redevelopment. The effects on renovation and deconversion permits are not statistically distinguishable from zero in either design, though the boundary design shows some degree of substitution toward renovation and a reduction in deconversion



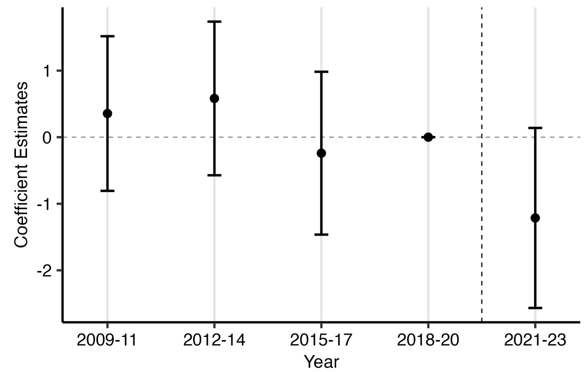
(a) Demolition Permits



(b) Construction Permits



(c) Renovation Permits



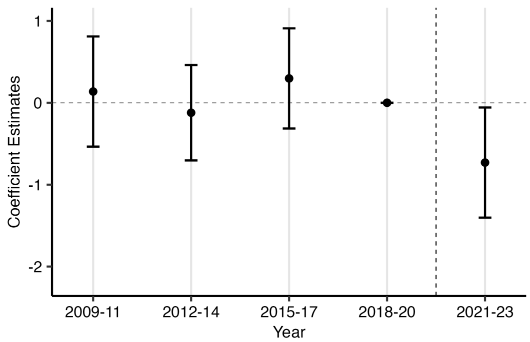
(d) Deconversion Permits

Figure D.4: Boundary Design: Building Permits

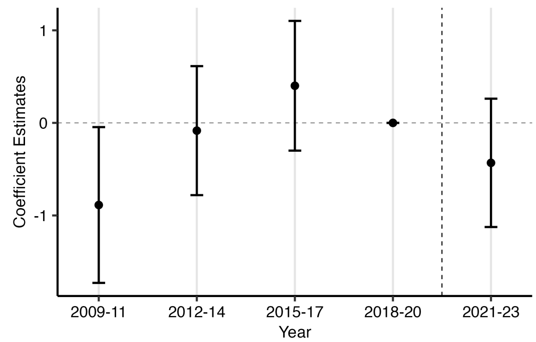
Note: Logit estimates of equation (16) on the boundary sample, three-year periods, 500m buffer, with parcel fixed effects, area-by-period fixed effects, and period-specific linear controls in longitude and latitude. Conley spatial standard errors with a 100-meter cutoff. Panels (a)–(d) report demolition, construction, renovation, and deconversion permits; renovation is an addition or remodel permit, and deconversion reduces the number of units on a parcel. Confidence intervals are at the 95% level.

permits.

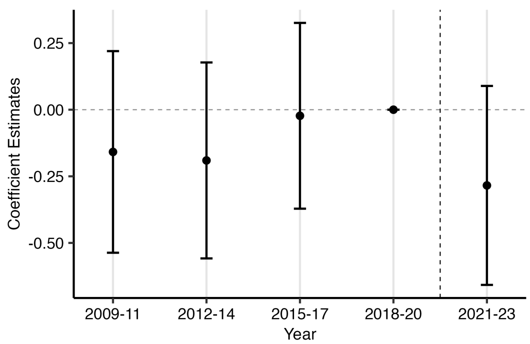
We also estimate the policies' effect on demolition separately for the two treated neighborhoods (Figure D.6). We find that the two treated neighborhoods respond similarly, at −0.88 in the 606 Trail and −1.53 in Pilsen with the difference between them insignificant, which supports treating σ_c as common across neighborhoods. Widening the ring to one kilometer leaves demolition and construction significant at the 1 percent level, and narrowing it to 250 meters leaves demolition at −0.95 with a standard error half again as large, too imprecise to detect (Table D.4). Permit processing times do not lengthen, so the effect is not administrative delay in another guise, and the estimated cost of the projects still being built does not move, so the decline is in the number of redevelopment projects rather than in their quality (Figure D.7).



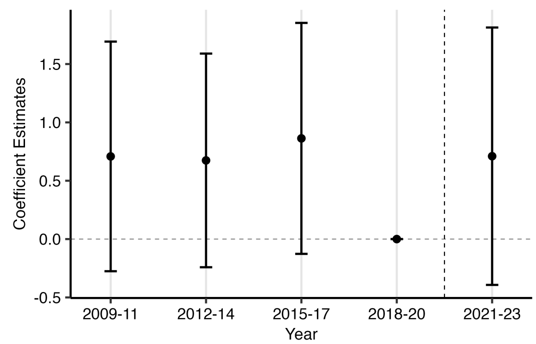
(a) Demolition Permits



(b) Construction Permits



(c) Renovation Permits



(d) Deconversion Permits

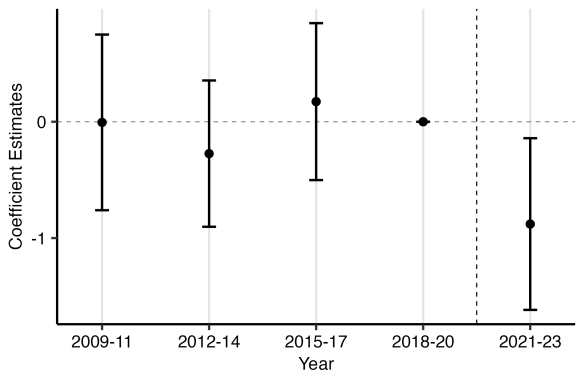
Figure D.5: Matched Design: Building Permits

Note: Logit estimates of equation (16) on the propensity-score matched sample of Table D.2, with period fixed effects in place of the area-by-period fixed effects. Panels (a)–(d) report demolition, construction, renovation, and deconversion permits, defined as in Figure D.4. Conley spatial standard errors with a 100-meter cutoff. The reference period is 2018–2020, and confidence intervals are at the 95% level.

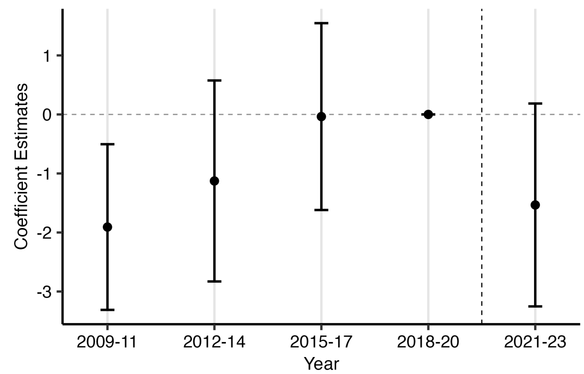
A logit with parcel fixed effects is identified from the conditional likelihood, which is degenerate for any parcel whose outcome never varies. The estimation therefore keeps only the 527 parcels in the boundary sample that are ever demolished, 3.8 percent of the total, and the coefficient describes when a parcel is redeveloped rather than whether it is. We re-estimate without them, keeping every parcel within 500 meters around the policy boundary. Table D.4 shows that the estimated effects on demolition and construction change little in either design, and each estimate lies inside the confidence interval of its fixed-effects counterpart.

D.4 Rents and Sale Prices

Neither rents nor sale prices respond (Figure D.8). Both estimates are negative and not distinguishable from zero throughout, before and after April 2021, and both are robust to the alternative buffer widths. Over three years, at an annual demolition rate below one percent, the housing the ordinances leave standing is a small addition to the stock, and



(a) 606 Trail



(b) Pilsen

Figure D.6: Boundary Design: Demolition Permits by Neighborhood

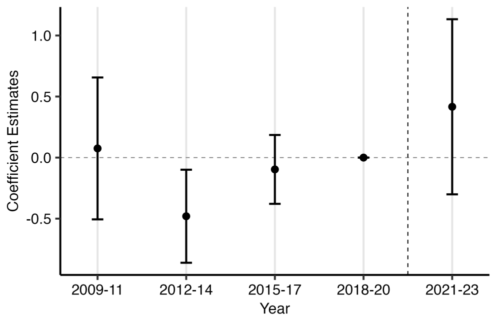
Note: The figure shows the logit estimation results of equation (16) at the three-year frequency with a 1km buffer, estimated separately for the 606 Trail and Pilsen neighborhoods. The dependent variable is the demolition permit dummy. Both regressions include parcel fixed effects, area-by-period fixed effects, and period-specific linear controls in longitude and latitude. Conley spatial standard errors with a 100-meter cutoff. Confidence intervals are at the 95% level.

the policy cannot significantly change either the income distribution or the amenities in the treated neighborhoods. Sale prices carry the sharper test, since they capitalize expectations. Anticipation, or an expectation that the ordinances would outlast April 2024, would have made the surcharge a lasting tax on redevelopment, eroding the option value embedded in older properties and depressing their prices. Prices instead trace a flat path through the policy date.

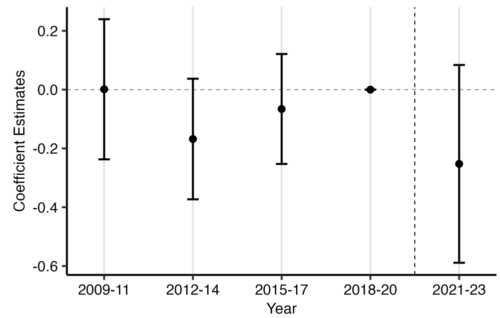
Table D.4: Demolition and Construction Permits: Buffer Width and Parcel Fixed Effects

	Demolition				Construction			
Sample	0.25km (1)	1km (2)	0.5km (3)	Matched (4)	0.25km (5)	1km (6)	0.5km (7)	Matched (8)
Treat ×								
2009–2011	0.014	−0.415	−0.321	0.177	0.956	0.013	0.379	−0.835**
	(0.564)	(0.340)	(0.369)	(0.283)	(0.927)	(0.449)	(0.574)	(0.356)
2012–2014	0.184	−0.460	−0.319	−0.096	0.490	0.297	0.454	−0.190
	(0.617)	(0.292)	(0.331)	(0.229)	(0.612)	(0.342)	(0.323)	(0.261)
2015–2017	0.654	0.001	0.133	0.186	0.754	0.204	0.338	0.157
	(0.536)	(0.301)	(0.334)	(0.234)	(0.486)	(0.304)	(0.276)	(0.243)
2018–2020	-	-	-	-	-	-	-	-
2021–2023	−0.946	−1.009***	−1.135***	−0.661**	0.141	−1.444***	−0.728**	−0.457*
	(0.678)	(0.335)	(0.422)	(0.292)	(0.576)	(0.342)	(0.328)	(0.271)
Parcel FE	Yes	Yes	No	No	Yes	Yes	No	No
Area × Period FE	Yes	Yes	Yes	No	Yes	Yes	Yes	No
$F_t(\text{Lon}, \text{Lat})$	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Num. parcels	293	994	13,726	24,743	256	883	13,726	24,743
Num. obs.	1,465	4,970	68,630	123,715	1,280	4,415	68,630	123,715
Mean Y	0.008	0.009	0.008	0.009	0.007	0.008	0.007	0.008

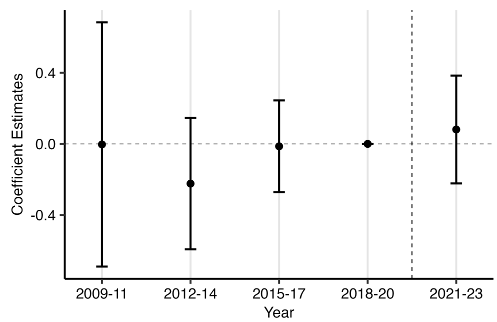
Note: The table shows the logit estimation results of equation (16): demolition permits in columns (1)–(4) and construction permits in columns (5)–(8), at the stated buffer width or on the matched sample. Columns with parcel fixed effects are identified only from parcels ever issued a permit of the given type; the others retain every parcel and add back the parcel-level main effects. All specifications include period-specific linear controls in longitude and latitude. Conley spatial standard errors with a 100-meter cutoff in parentheses. The reference period is 2018–2020. * p<0.1, ** p<0.05, *** p<0.01.



(a) Demolition Permits



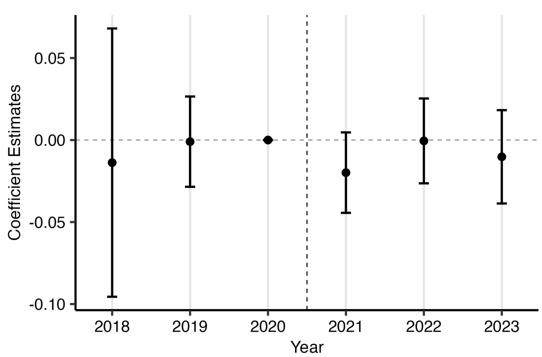
(b) All Permits



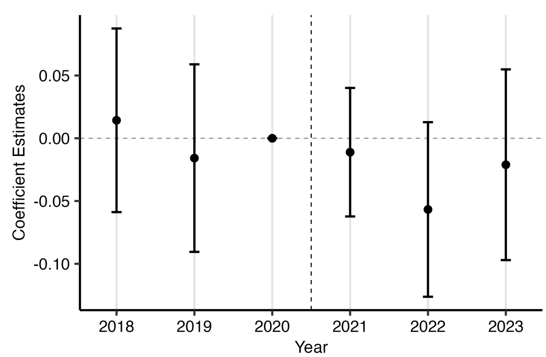
(c) Construction Cost

Figure D.7: Boundary Design: Permit Processing Time and Construction Cost

Note: The figure shows the OLS estimation results of equation (16) at the three-year frequency with a 500m buffer for log permit processing time (measured in days) and log estimated project costs of construction permits. All specifications include area-by-period fixed effects and period-specific linear controls in longitude and latitude. Conley spatial standard errors use a 100-meter cutoff. Confidence intervals are at the 95% level.



(a) Rental Prices



(b) Sales Prices

Figure D.8: Boundary Design: Rental and Sale Prices

Note: Estimates of equation (16), 500m buffer, for log rental price (Panel (a)) and log sales price (Panel (b)). The rent regression includes unit fixed effects; the price regression controls for housing characteristics. Confidence intervals are at the 95% level.

E Counterfactual Appendix

E.1 Equivalent Variation

We measure the welfare effect of a policy on a household by its equivalent variation: the increase in income, received in every location, that makes the household as well off at baseline rents and amenities as it is under the policy. Let superscript 0 denote the baseline economy and superscript 1 the counterfactual, and consider a household of type (z, x_0) in period t.

Household welfare. Given the rent function $P_t^s(q, x)$ in scenario $s \in \{0, 1\}$, a household with income m to spend obtains utility

$$U_t^s(x, m) = \max_q \, (q - \bar{q})^\alpha \left(m - P_t^s(q, x)\right)^{1-\alpha} \tag{E.1}$$

in neighborhood x, which is the housing quality problem (2)–(3) with income m in place of z and the numeraire eliminated through the budget constraint; it is zero if the household cannot afford any quality above $\bar{q}$ in x. Because the idiosyncratic preference shocks are i.i.d. Type-II extreme value with dispersion σ_x and the neighborhood choice is static, the expected utility of a type-(z, x_0) household from its neighborhood choice, before it observes its shocks, equals $\Gamma(1 - 1/\sigma_x)\, W_t^s(m; z, x_0)$, where

$$W_t^s(m; z, x_0) = \left[\sum_{x \in \mathbb{X}} \Big(\tau(x; x_0)\, U_t^s(x, m)\, A_t^s(x, z)\Big)^{\sigma_x} + \tau(o; x_0)^{\sigma_x}\right]^{1/\sigma_x} \tag{E.2}$$

is the inclusive value of its choice set, $A_t^s(x, z) = \bar{A}(x, z)\left[\bar{z}_t^{\,s}(x)\right]^\eta$ is the amenity in (6) under scenario s, and the constant $\Gamma(1 - 1/\sigma_x)$ is the same in both scenarios and drops out of every welfare comparison. Income m enters the utility of every neighborhood through the budget constraint, whereas the household's type z, which determines the amenities it enjoys, and its origin x_0, which determines its moving costs, are held fixed; the outside option retains its normalized utility of one.

Equivalent variation. The equivalent variation of the policy for a type-(z, x_0) household in period t, $\mathrm{EV}_t(z, x_0)$, is the increase in income that raises its welfare in the baseline economy to the level it attains under the policy:

$$W_t^0\big(z + \mathrm{EV}_t(z, x_0);\; z, x_0\big) = W_t^1\big(z;\; z, x_0\big). \tag{E.3}$$

The right-hand side is the household's welfare under the policy, when it faces counterfactual rents and amenities at its own income. The left-hand side is its welfare at baseline rents and amenities when it receives $\mathrm{EV}_t(z, x_0)$ on top of its income. Because the increase in income is received wherever the household lives, it raises $U_t^0(x, z + \mathrm{EV}_t)$ in every neighborhood at once, and the household reoptimizes its neighborhood and its housing quality when compensated. The equivalent variation is positive if the household gains from the policy and negative if it loses, and we express it as a share of income, $\mathrm{EV}_t(z, x_0)/z$. Since W_t^0 is continuous and strictly increasing in income for any household that can afford housing in at least one neighborhood, equation (E.3) has a unique solution.

F Additional Figures and Tables

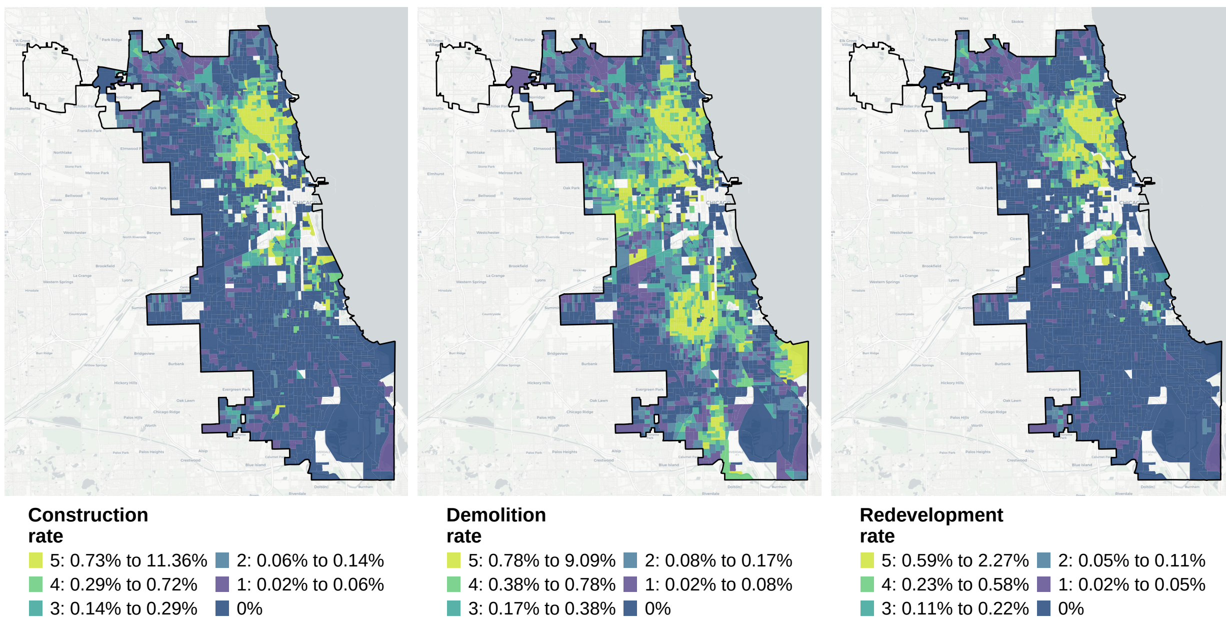


Figure F.1: Construction, Demolition, and Redevelopment Rates across Chicago Block Groups

Note: The figure maps, from left to right, the construction, demolition, and redevelopment rates across Chicago block groups, each defined as the share of 2005 parcels receiving a permit of that type in a year, averaged over 2009–2019; a parcel counts as redeveloped only when its demolition permit is matched to a new-construction permit on the same parcel. Block groups with no such permit form their own category, shaded dark blue, and the rest are split into quintiles of the positive rates. Data source: Building Permit Data from the City of Chicago Data Portal.

Table F.1: Permit rates on block group characteristics: non-linear specification

Dependent variable:	Permit rate (%)							
	Full sample				Excluding extreme teardowns			
	Demo.	Constr.	Redev.	Renov.	Demo.	Constr.	Redev.	Renov.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log household income	-6.76^{***}	0.94^{**}	0.49	-2.88^{***}	0.96	1.24	1.19^{*}	-2.34^{**}
	(0.69)	(0.42)	(0.35)	(0.67)	(0.81)	(0.86)	(0.70)	(1.10)
squared	0.30^{***}	-0.05^{**}	-0.02	0.12^{***}	-0.05	-0.06	-0.06^{*}	0.10^{*}
	(0.03)	(0.02)	(0.02)	(0.03)	(0.04)	(0.04)	(0.03)	(0.05)
Log gross rent	1.56	-2.03^{***}	-1.53^{***}	-2.01^{**}	-1.77^{*}	-2.53^{**}	-1.84^{**}	-3.27^{**}
	(1.00)	(0.61)	(0.50)	(0.98)	(0.95)	(1.01)	(0.83)	(1.29)
squared	-0.10	0.15^{***}	0.12^{***}	0.16^{**}	0.14^{*}	0.19^{***}	0.14^{**}	0.24^{***}
	(0.07)	(0.04)	(0.04)	(0.07)	(0.07)	(0.07)	(0.06)	(0.09)
Log home value	-14.91^{***}	-8.34^{***}	-7.78^{***}	-2.17^{***}	-13.48^{***}	-11.73^{***}	-11.36^{***}	-3.76^{***}
	(0.75)	(0.45)	(0.38)	(0.73)	(0.84)	(0.89)	(0.72)	(1.14)
squared	0.61^{***}	0.35^{***}	0.33^{***}	0.11^{***}	0.56^{***}	0.49^{***}	0.47^{***}	0.17^{***}
	(0.03)	(0.02)	(0.02)	(0.03)	(0.03)	(0.04)	(0.03)	(0.05)
Building age (decades)	-0.20^{***}	-0.18^{***}	-0.13^{***}	-0.43^{***}	-0.14^{***}	-0.19^{***}	-0.13^{***}	-0.32^{***}
	(0.06)	(0.03)	(0.03)	(0.06)	(0.05)	(0.05)	(0.04)	(0.06)
squared	0.02^{***}	0.01^{***}	0.01^{***}	0.04^{***}	0.01^{***}	0.02^{***}	0.01^{***}	0.03^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)
Observations	2,331	2,331	2,331	2,331	1,358	1,358	1,358	1,358
R^2	0.33	0.43	0.46	0.32	0.48	0.43	0.48	0.36

Notes: This table replicates Table 1, adding squared terms in log income, log gross rent, and log home value. Columns (1)–(4) use the full sample; columns (5)–(8) exclude block groups whose demolition rate exceeds 1.5 times their construction rate. "Redev." is the redevelopment rate defined in Table 1. All right-hand side variables are block-group medians from the ACS five-year estimates. Permit rates are shares of the number of parcels in 2005 with a permit of the corresponding type, averaged over 2009–2019. Block-group level characteristics are calculated using the average five-year ACS data from 2009–2019. Regressions weighted by number of parcels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table F.2: Comparing characteristics of newly developed and redeveloped housing across neighborhoods.

Dependent variable	New Building			Change from Old to New		
	Units	ln (Total SF)	ln (SF/unit)	Units	Δ ln (Total SF)	Δ ln (SF/unit)
Log income	-0.32^{***}	0.06	0.21^{***}	-0.22	0.10	0.19^{***}
	(0.08)	(0.05)	(0.05)	(0.13)	(0.06)	(0.05)
Log gross rent	0.18	-0.01	-0.10	-0.17	-0.13^{*}	-0.01
	(0.11)	(0.07)	(0.07)	(0.12)	(0.07)	(0.08)
Log housing value	-0.05	0.24^{***}	0.26^{***}	-0.35^{***}	-0.03	0.18^{***}
	(0.09)	(0.07)	(0.06)	(0.09)	(0.06)	(0.05)
Building age (decades)	0.01	0.18	0.16	0.08	0.13	0.09
	(0.08)	(0.11)	(0.11)	(0.14)	(0.09)	(0.06)
Building age squared	0.00	-0.02	-0.02	-0.01	-0.01	-0.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Mean dep. var.	1.36	7.85	7.69	-0.46	0.56	0.86
Observations	6,283	6,247	6,247	2,715	2,698	2,698
R^2	0.07	0.02	0.04	0.08	0.02	0.05

Notes: All right-hand side variables are block-group medians from the ACS five-year estimates. All regressions include year fixed effects and are unweighted. Standard errors are clustered at the Chicago official neighborhood level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table F.3: First stage: redevelopment rate on commuting shock × share built before 1910

Dependent variable	Redevelopment rate
	(1)
Commuting shock × share built before 1910	62.81^{***}
	(12.37)
Commuting shock	Yes
Initial log income	Yes
Vintage share controls	Yes
Cubic in distance to the Loop	Yes
Observations	1,277
R^2	0.39
Kleibergen-Paap F-stat	25.80

Notes: First stage of the IV estimates in Table 2, regressing the redevelopment rate on the instrument, the interaction of the commuting shift-share labor demand shock described in the text with the initial (2009) share of housing built before 1910. Sample, controls, weighting, and clustering are as in Table 2. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table F.4: The effects of redevelopment on neighborhood outcomes: robustness

	(1)				(2)		(3)		(4)	
	Value	Inc	Services	Crime	Value	Inc	Value	Inc	Value	Inc
Panel A: OLS, minimum of demolition and construction rates										
Redevelopment rate	0.02***	0.08***	0.02***	-0.01**	0.02***	0.06***	0.01***	0.04***	—	—
	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)		
Panel B: IV, minimum of demolition and construction rates										
Redevelopment rate	0.10***	0.16***	0.03	-0.09**	0.12***	0.18***	0.09**	0.12***	0.09**	0.14***
	(0.04)	(0.04)	(0.05)	(0.04)	(0.04)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)
Kleibergen-Paap F	25.8				26.0		19.5		24.2	
Panel C: IV, demolition rate										
Demolition rate	0.03**	0.05***	0.01	-0.03***	0.04**	0.06***	0.04**	0.05***	0.03**	0.04***
	(0.01)	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.02)
Kleibergen-Paap F	43.1				53.0		29.5		65.9	
Panel D: IV, linked demolition and construction permit rate										
Redevelopment rate	0.10***	0.17***	0.03	-0.10**	0.13**	0.20***	0.12**	0.15**	0.09**	0.15***
	(0.04)	(0.05)	(0.05)	(0.04)	(0.05)	(0.06)	(0.05)	(0.06)	(0.04)	(0.05)
Kleibergen-Paap F	24.6				24.1		9.8		23.9	
Sample	SFq > 0.8				SFq > 0.6		All BGs		SFq > 0.8	
Observations	1,277				1,726		2,223		1,277	
Full controls	Y				Y		Y		Y	
Instrument	Pre-1910				Pre-1910		Pre-1910		Pre-1920	

Notes: Estimates of equation (1): 2009–2019 log changes in median housing value (Value) and median income (Inc). Column (1) additionally reports the log change in the number of consumer-services establishments (Services; NAICS 44–45, 71, 72 and 81) and in the number of reported crime incidents (Crime), both as three-year averages over 2008–2010 and 2017–2019, which smooths year-to-year fluctuation in establishment coverage; establishment counts are floored at one before taking logs. Both series are described in Section 2. Columns: (1) block groups with initial (2009) share of units in 1–4 unit buildings (SFq) above 0.8, as in Table 2; (2) SFq above 0.6; (3) all block groups; (4) as (1), with the pre-1920 share in the instrument. Panels: A, OLS on the minimum of the demolition and construction rates; B, IV with the same regressor; C, IV with the demolition rate; D, IV with the rate of demolition permits linked to a construction permit on the same parcel. Controls, weights, and clustering are as in Table 2. OLS is identical in columns (1) and (4) and is shown once. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table F.5: Neighborhood Groups

Group	Neighborhoods
1	Douglas, Grand Boulevard, Hyde Park, Kenwood, Oakland
2	Loop, Old Town, Printers Row, River North, Streeterville
3	Garfield Park, Little Italy, UIC, North Lawndale, United Center
4	Bucktown, Lincoln Park, North Center, Sheffield & DePaul
5	Austin, Belmont Cragin, Hermosa, Humboldt Park
6	Archer Heights, Brighton Park, Gage Park, Little Village, New City
7	Armour Square, Bridgeport, Chinatown, Lower West Side, McKinley Park
8	Avalon Park, Calumet Heights, South Chicago
9	Burnside, East Side, Hegewisch, Pullman, Riverdale, South Deering
10	Englewood, Fuller Park, Grand Crossing, Washington Park
11	Jackson Park, South Shore, Woodlawn
12	Dunning, O'Hare, Portage Park
13	Albany Park, Avondale, Irving Park, Logan Square
14	Andersonville, Edgewater, Lincoln Square, North Park, Rogers Park, Uptown, West Ridge
15	Edison Park, Jefferson Park, Norwood Park, Sauganash, Forest Glen
16	Ashburn, Auburn Gresham, Chatham, Chicago Lawn
17	Morgan Park, Roseland, Washington Heights, West Pullman
18	Clearing, Garfield Ridge, West Elsdon, West Lawn
19	Beverly, Mount Greenwood
20	Greektown, Near South Side, West Loop
22	Boystown, Lake View, Wrigleyville
24	East Village, Ukrainian Village, West Town, Wicker Park
25	Galewood, Montclare

Note: Each row lists the Chicago community areas forming one model neighborhood group. The table lists all 23 groups the model solves. Numbering follows the source data, in which groups 21 and 23 do not appear: the community areas that would compose them are dominated by high-rise apartment buildings outside the one-to-six-unit scope of the model and are excluded from the sample.

Table F.6: Hedonic Pricing Function Estimates

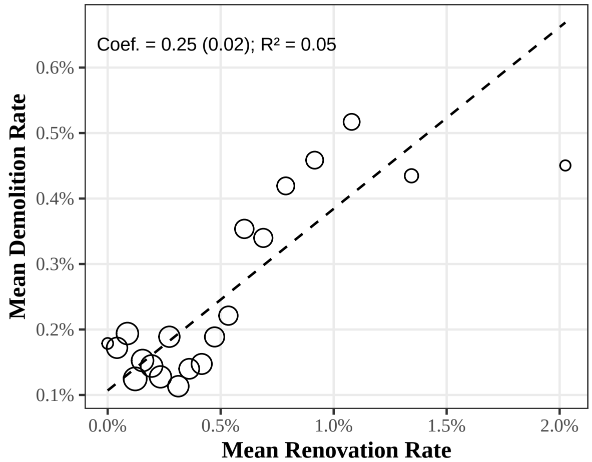
	Coeff.		Coeff.
Depreciation & Renovation		*Heating (omit: Forced Air)*	
Building Age (years)	−0.0034	Electric	0.1443
Renovation Indicator	0.2343	Hot Water/Steam	0.0306
		None	−0.0545
Unit Characteristics		*Building Type (omit: 2 Story)*	
Bathrooms per Unit	−0.0716	1.5 Story	0.1727
Bathrooms per Unit^2	0.0132	1 Story	0.1407
Bedrooms per Unit	0.0298	3+ Story	0.0245
Bedrooms per Unit^2	0.0010	Split Level	0.1134
Log Land Area per Unit	−0.0419		
Log Floor Area per Unit	0.2265		
Constr. Quality (omit: Average)		*Other Features*	
Deluxe	0.2990	Garage	0.0123
Poor	−0.1150	Frame Enc. Porch	0.0336
		Masonry Enc. Porch	0.0556
Exterior Wall (omit: Masonry)		*Property Type (omit: Small MF)*	
Frame	0.1024	Medium Multifamily	−0.0251
Frame & Masonry	0.1228	Single Family	0.1961
Stucco	−0.0186		
Neighborhood FE: Yes		Observations: 1,263,684	
Neighborhood Slope: Yes		R^2: 0.94	

Note: Renormalized coefficients from the nonlinear hedonic regression $\log P_{it} = \log \kappa(x) + \nu(x) \log q_{it}$, with $\log q_{it} = -\delta \, \text{age}_{it} + \sum_{c \in \mathcal{C}} \alpha_c X_{it}^c + \xi_{it}$. Land and floor area per unit enter in logs; other characteristics in levels or as indicators. Neighborhood rent levels $\kappa(x)$ and quality elasticities $\nu(x)$ are estimated jointly and reported in Figure F.5. Sample: properties built after 1930, 2014–2020, demeaned by year.

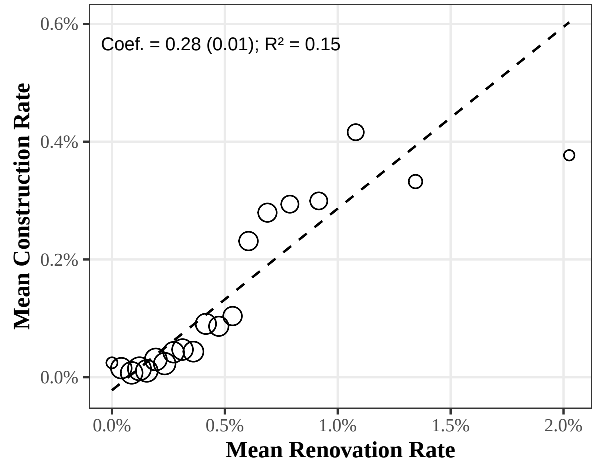
Table F.7: Pricing Function Correlates across Neighborhoods

	Elasticity $\nu(x)$ (1)	Log Level $\log\kappa(x)$ (2)
Log(Median Building Age)	0.250^{*}	-0.193
	(0.133)	(0.161)
Log(Average Income)	0.258^{**}	0.080
	(0.100)	(0.121)
Constant	-3.010^{**}	6.095^{***}
	(1.421)	(1.716)
Observations	23	23
R^2	0.283	0.116

Note: Each column regresses a neighborhood-group-level rent function estimate on log median building age and log average household income, weighted by housing units. Both columns use the estimates from the nonlinear hedonic regression (15): $\nu(x)$ is the quality elasticity of rent and $\kappa(x)$ is the rent level. Standard errors in parentheses. *p<0.1; **p<0.05; ***p<0.01.



(a) Renovation and demolition



(b) Renovation and construction

Figure F.2: Renovation, demolition, and construction across neighborhoods in Chicago, 2009–2019

Notes: Each panel presents a binscatter of two permit rates across Chicago block groups, 2009–2019. Block groups are ranked by the renovation rate and divided into 20 equal-frequency bins; dots plot the parcel-count-weighted mean of each rate within a bin, with dot size proportional to the total number of parcels in the bin. Permit rates are computed as shares of the number of parcels in 2005. Regression coefficients, standard errors, and R^2 are from weighted OLS on the underlying block-group-level data, with weights equal to the number of parcels. Block groups with a mean renovation rate above 10% are excluded.

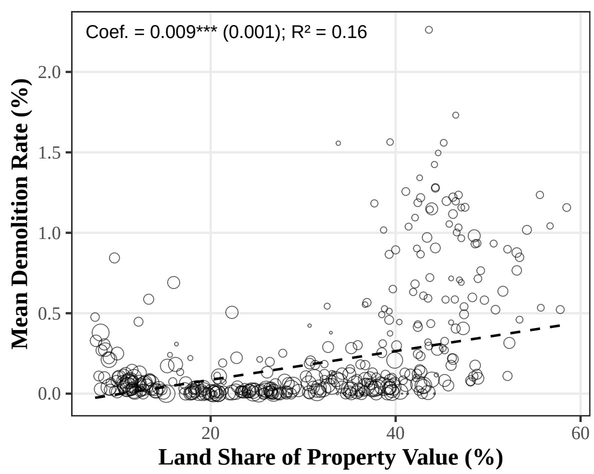


Figure F.3: Land share and demolition rate across census tracts

Notes: Each point is a census tract within the City of Chicago. The x-axis is the land share of housing value from Davis et al. (2021); the y-axis is the mean demolition permit rate. Both the cross-sectional land share estimates and the mean demolition rate are calculated over 2012–2022. The positive relationship indicates that demolition activity is more intensive in tracts where land accounts for a larger share of property value—i.e., where structures have low residual value relative to their sites.

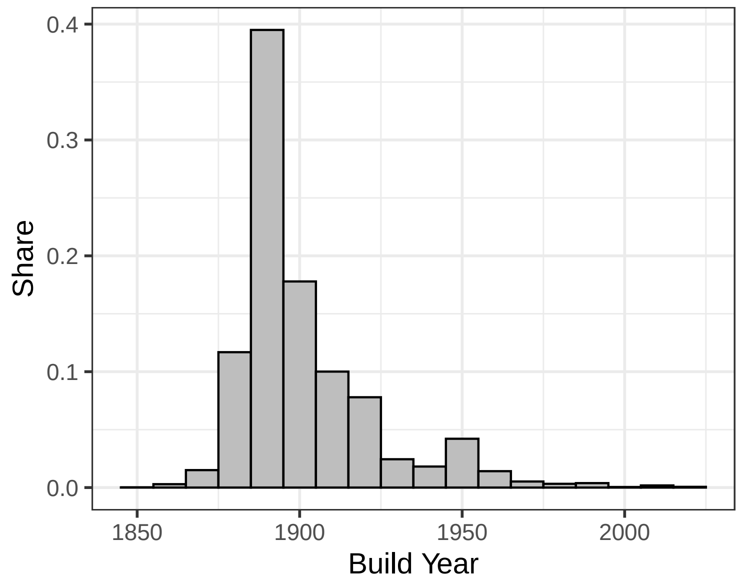


Figure F.4: Distribution of Build Year of Redeveloped Buildings

Notes: The figure plots the distribution of the build year of the structure that was replaced, across redevelopment events in Chicago. Bars are ten-year bins and heights are shares of all plotted events, so they sum to one. A redevelopment event is an assessment build-year jump on a parcel, as described in Section 3.2; events matched only to a nearby renovation permit, with no nearby construction permit, are excluded because the build-year rule alone cannot distinguish a major renovation from a rebuild. The sample is the 6,504 Chicago events for which the pre-redevelopment build year is recorded in the assessment data. The median replaced structure was built in 1893, and 76 percent were built before 1910.

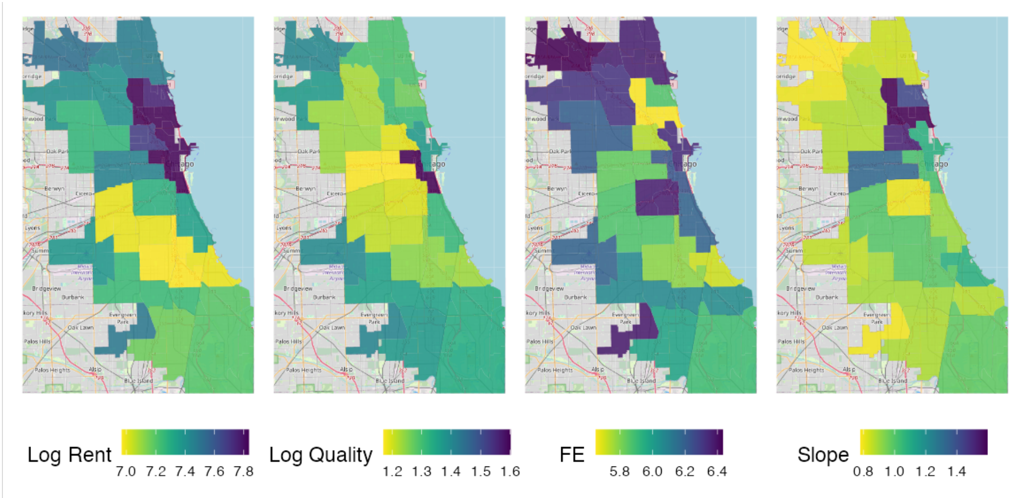
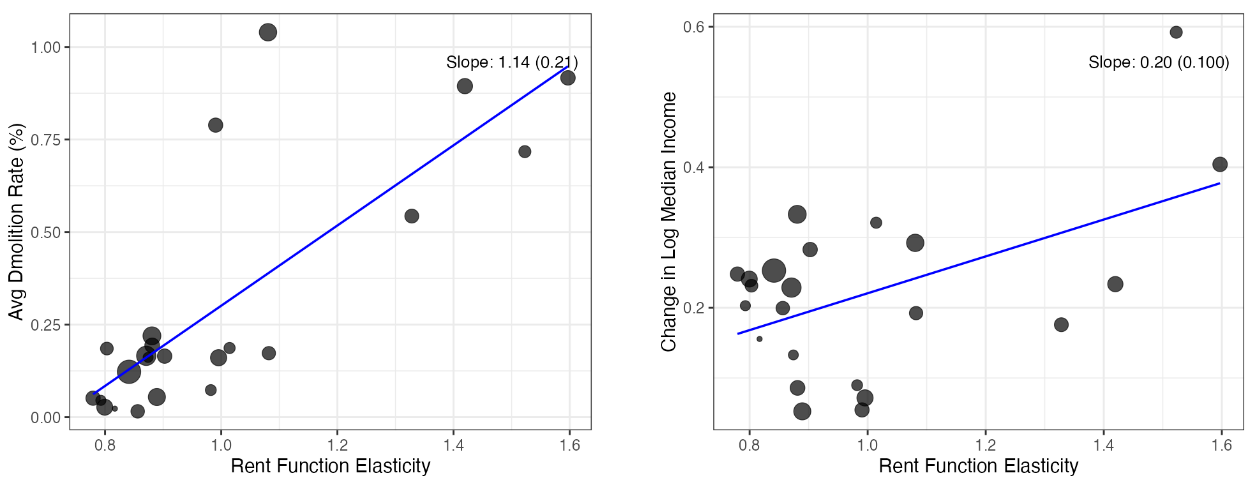


Figure F.5: Housing rent, quality and estimated rent functions across neighborhoods



(a) Demolition rate

(b) Income growth

Figure F.6: Rent function elasticity $\nu(x)$ and neighborhood dynamics

Notes: Each point is a neighborhood group, sized by housing units. Panel (a) plots the average annual demolition rate against the estimated rent function elasticity $\nu(x)$. Panel (b) plots the 2010–2019 change in log median household income against $\nu(x)$. Both regressions are weighted by housing units.

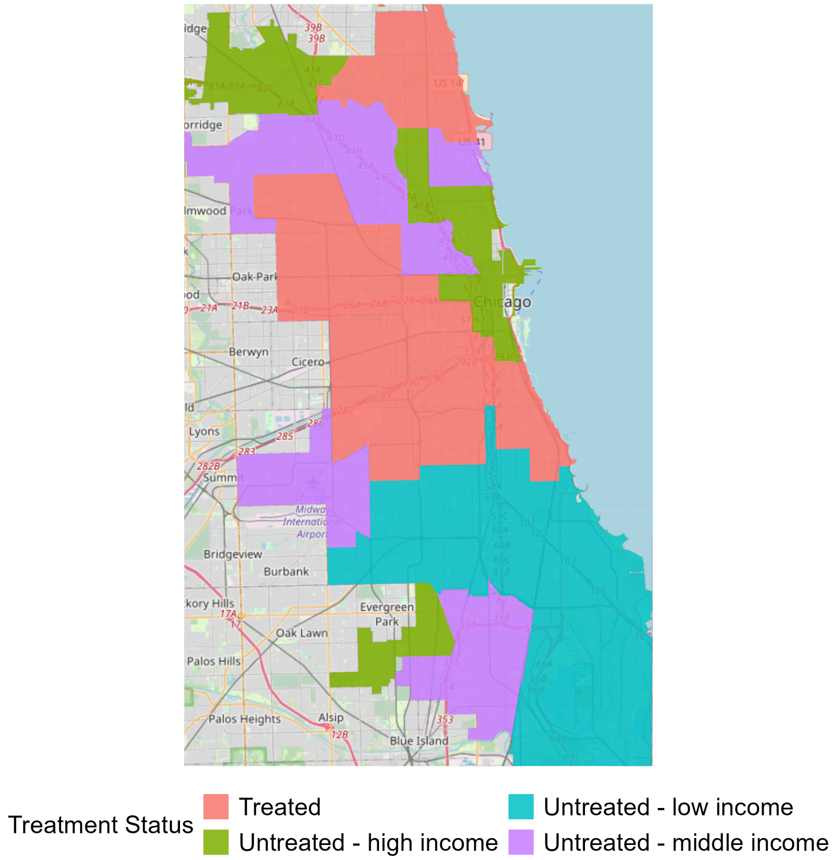
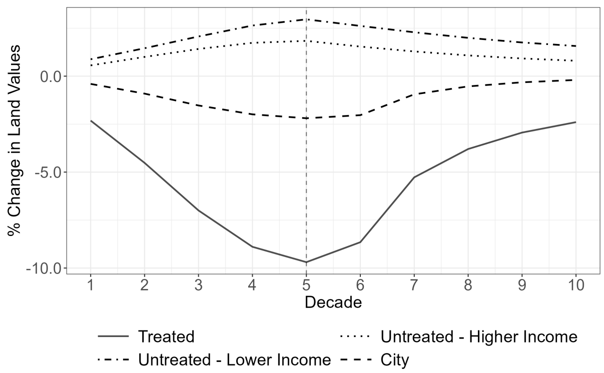
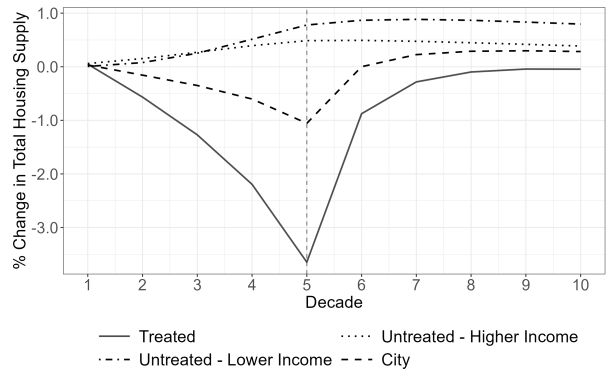


Figure F.7: Map of neighborhood groups by treatment status for the Anti-Redevelopment Policy

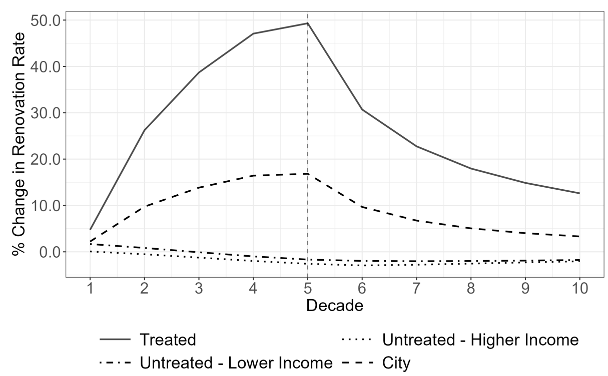
Note: Treatment area represents 60% of the total population of the Chicago municipality (approximately 700,000 housing units). Untreated areas take the remaining share.



(a) Land values



(b) Housing units



(c) Renovation rate

Figure F.8: The Impact of the Anti-Redevelopment Policy across Neighborhoods

Notes: Each panel plots the percent deviation from the baseline economy, by decade, for the treated neighborhoods, the untreated neighborhoods split at the median of average income, and the city as a whole. The policy is in force for the first five decades; the vertical dashed line marks its expiry. Panel (a) reports land values, Panel (b) the total number of habitable housing units, and Panel (c) the renovation rate.

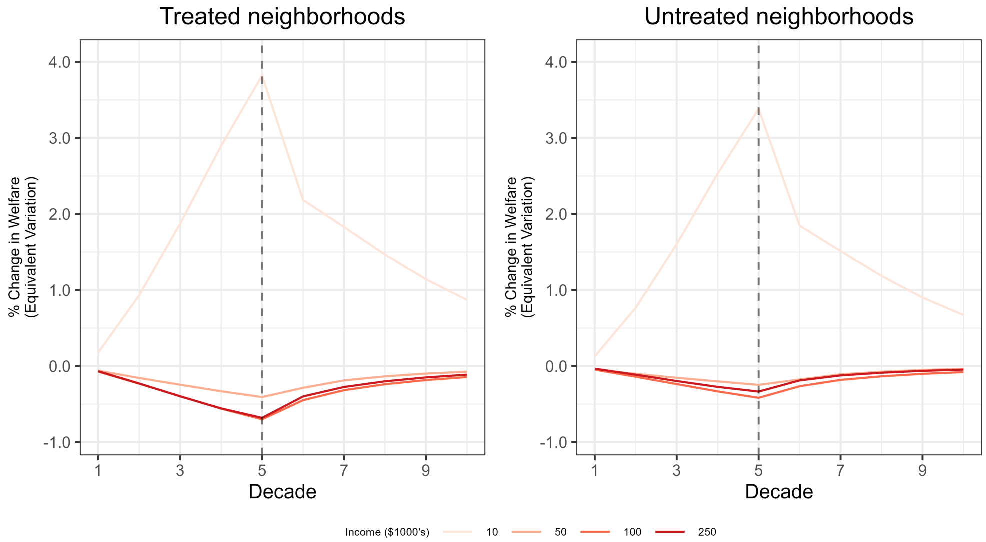


Figure F.9: The Impact of the Anti-Redevelopment Policy on Welfare, by Household's Initial Neighborhood

Note: Percent change in welfare (equivalent variation) by decade, for households earning $10,000, $50,000, $100,000, and $250,000, separately for those originally living in the treated neighborhoods (left) and in the untreated neighborhoods (right). The policy is in force for the first five decades; the vertical dashed line marks its expiry.

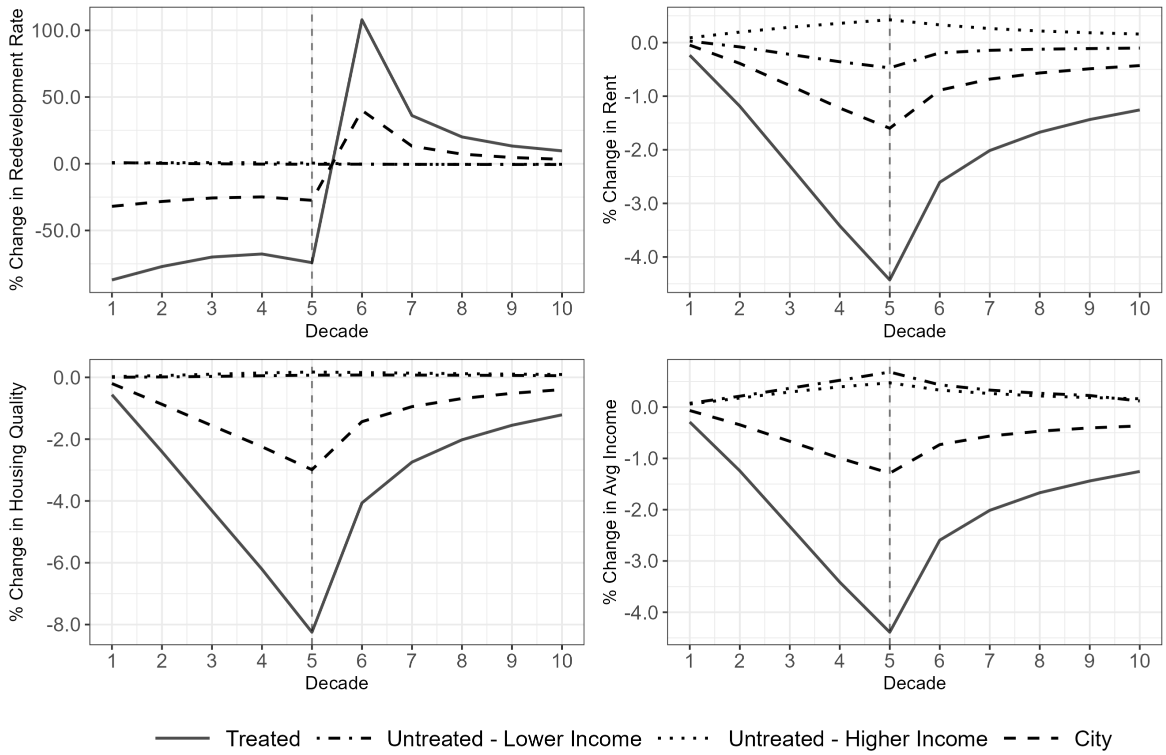


Figure F.10: The Impact of the Anti-Redevelopment Policy without Endogenous Amenities on Housing Redevelopment, Rent, Quality and Neighborhood Income

Note: Counterpart of Figure 5 in which neighborhood amenities are held fixed at their baseline values. Percent change relative to the baseline economy in the same decade. The vertical line marks the last decade in which the policy is in force.

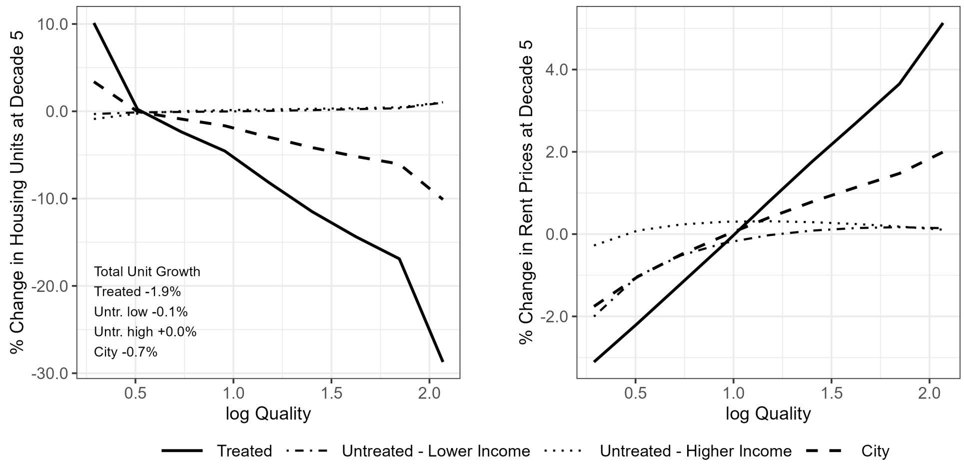


Figure F.11: The Impact of the Anti-Redevelopment Policy without Endogenous Amenities on Units and Rent by Quality

Note: Counterpart of Figure 6 in which neighborhood amenities are held fixed at their baseline values. Percent change in the number of housing units (left) and in the rent function (right) against log housing quality, in decade 5, relative to the baseline economy. The inset reports the percent change in the total number of units for each neighborhood group. We exclude from the unit counts units that have depreciated to the minimum quality $\bar{q}$ and are no longer habitable.

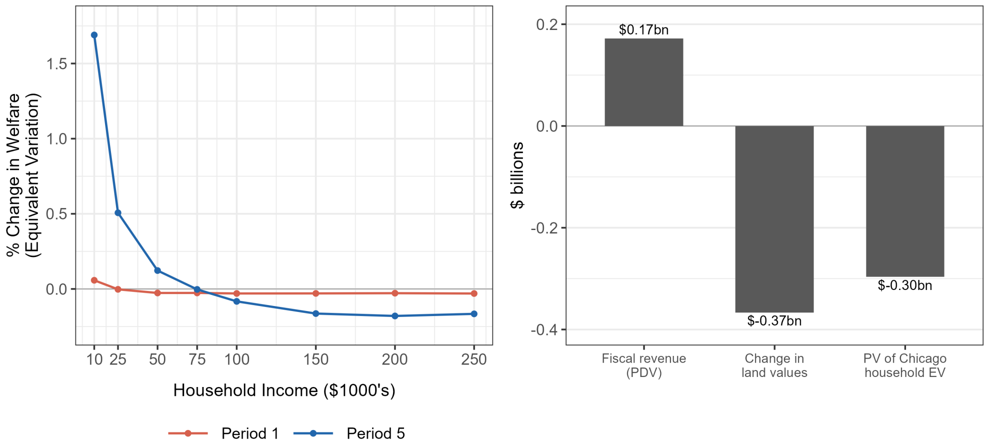


Figure F.12: The Impact of the Anti-Redevelopment Policy without Endogenous Amenities on Welfare and City Aggregates

Note: Counterpart of Figure 7 in which neighborhood amenities are held fixed at their baseline values. The left panel reports the equivalent variation as a share of income, by household income, in the first and last decade of the policy. The right panel reports the changes in the discounted values of fiscal revenue and of household equivalent variation over the model horizon, as well as the change in land values in period 1 of the policy. The present discounted value of household equivalent variation is calculated only for households whose origin is within the Chicago municipality.

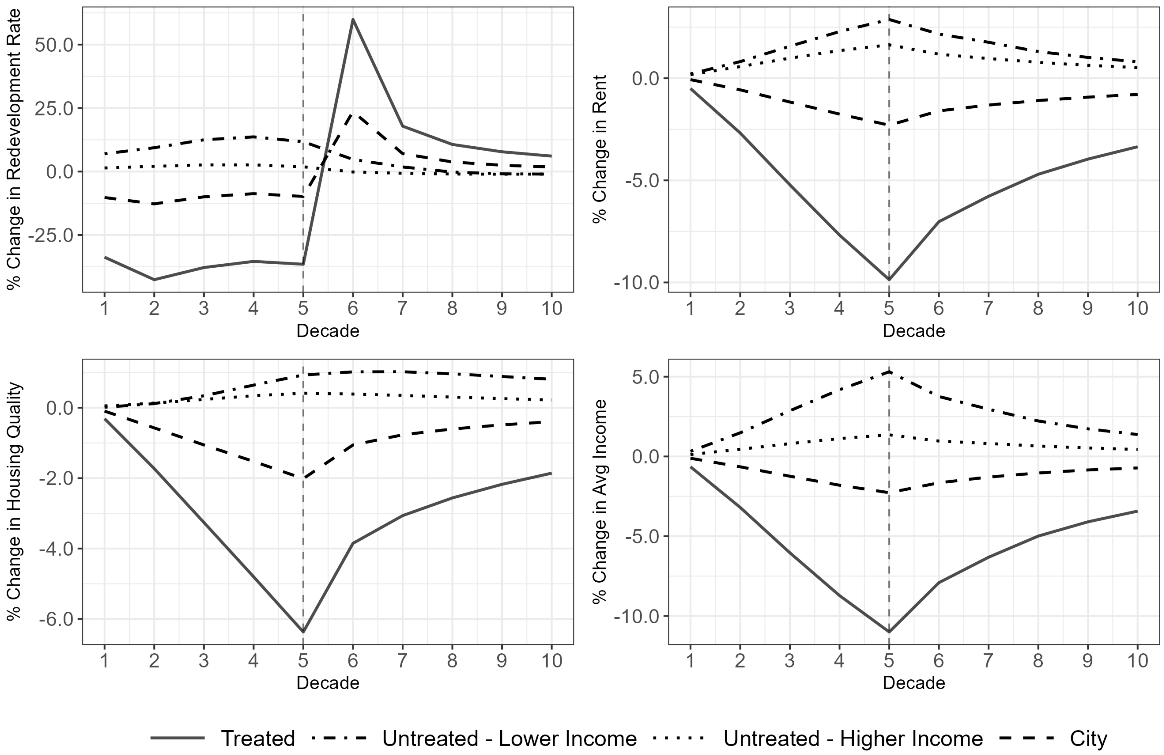


Figure F.13: The Impact of the Anti-Deconversion Ordinance on Housing Redevelopment, Rent, Quality and Neighborhood Income

Note: Counterpart of Figure 5 where only the anti-deconversion ordinance is imposed without the accompanying teardown tax. Percent change relative to the baseline economy in the same decade. The vertical line marks the last decade in which the policy is in force.

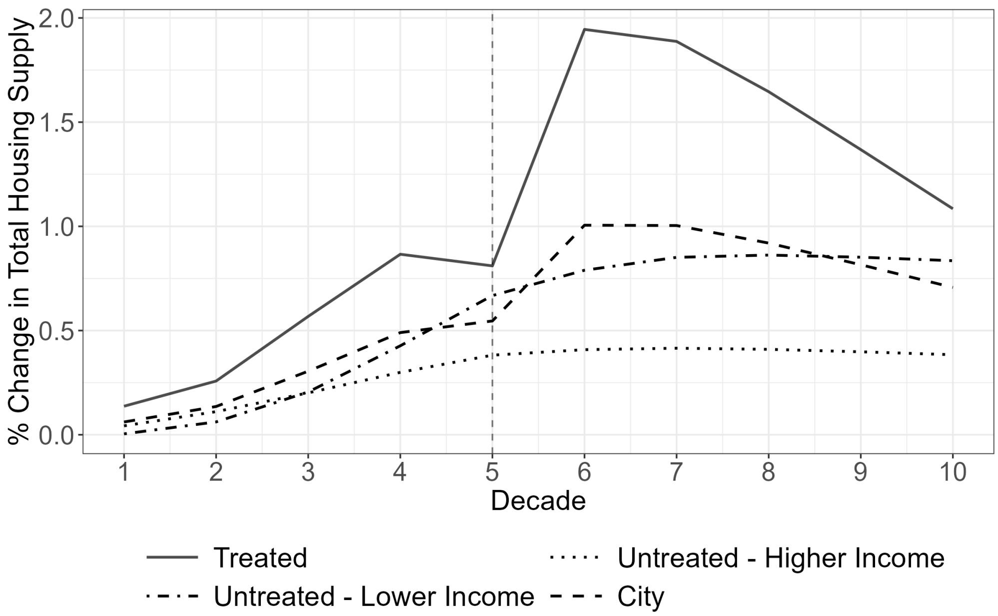


Figure F.14: The Impact of the Anti-Deconversion Ordinance on the Size of the Housing Stock

Note: Counterpart of Figure F.8b where only the anti-deconversion ordinance is imposed without the accompanying teardown tax. Percent change relative to the baseline economy in the same decade. The vertical line marks the last decade in which the policy is in force.

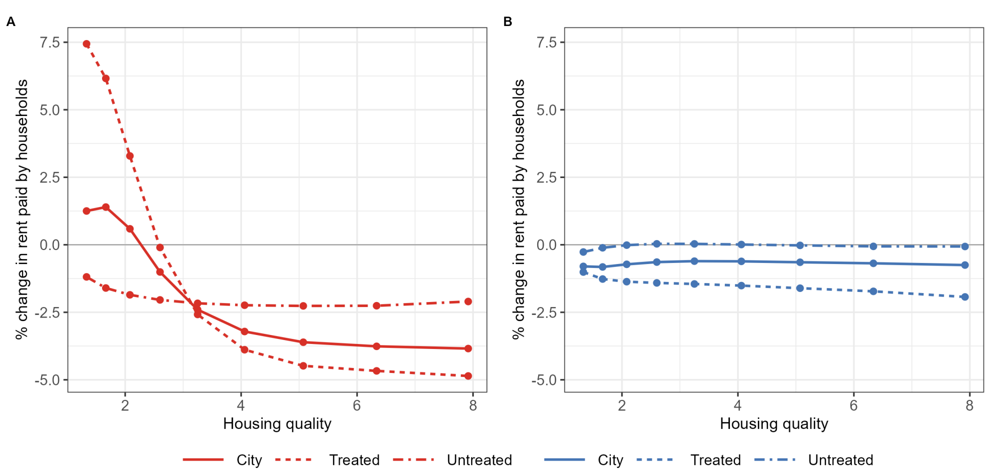


Figure F.15: Spillover effects of Subsidies on Rent

Note: Counterpart of Figure 8, where rent changes across the quality distribution are broken down by neighborhood treatment status. Red denotes the redevelopment subsidy, while Blue denotes the rent subsidy to demand. Redevelopment subsidies cause untreated neighborhoods to decrease in income because they incentivize the out-migration of high income households.

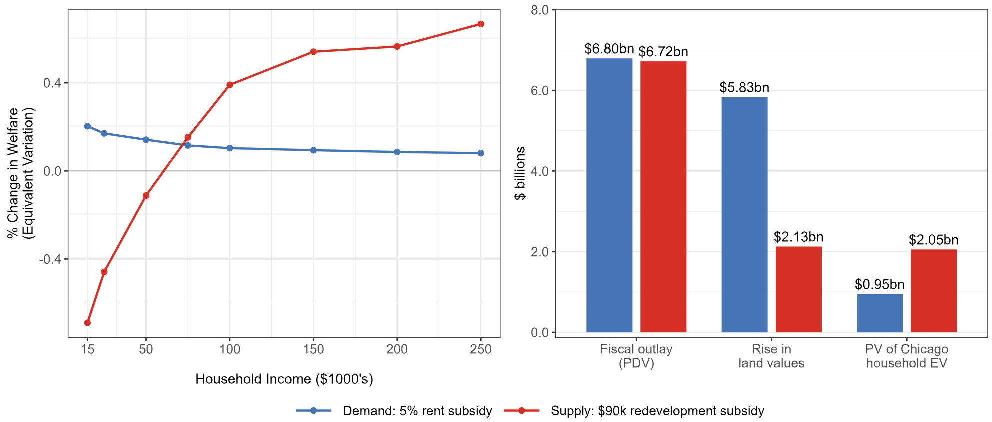


Figure F.16: Welfare and Fiscal Incidence of the Redevelopment and Rent Subsidies without Endogenous Amenities

Note: Counterpart of Figure 9 in which neighborhood amenities are held fixed at their baseline values. The left panel reports the equivalent variation as a share of income in decade 5, by household income. The right panel reports the period-1 present discounted values of the fiscal outlay over the policy window, the change in aggregate land values, and household equivalent variation over the model horizon. The present discounted value of household equivalent variation is calculated only for households whose origin is within the Chicago municipality.