Land use regulation, homeownership, and the cost of NIMBYism*

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Abstract — We examine how homeownership influences land use regulation and, consequently, impacts welfare and spatial sorting. Utilising panel data from England and employing the Right-to-Buy policy as an exogenous shifter in local homeownership, we first establish that a higher homeownership share is associated with (i) a higher share of refused planning applications, (ii) reduced construction, (iii) a lower share of developed land, and (iv) higher house prices, all supporting the 'homevoter hypothesis'. Building on these findings, we set up a dynamic spatial equilibrium model in which increased regulatory restrictiveness results in decreased construction activity but possibly elevated amenity levels. We conduct counterfactual analyses by eliminating the political influence of local homeownership to evaluate the welfare costs of NIMBYism.

Keywords — land use regulation, homeownership, NIMBYism housing supply, land, sorting, spatial inequality.

JEL codes — H43, R21, R30

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1 Introduction

The issue of reduced housing affordability is a global concern amid increasing demand for urban living. Over the past decades, real property prices in London have more than doubled. This trend is not unique to London; smaller cities like Birmingham, Manchester, and Bath also witnessed strong price appreciation, which have not been paralleled by comparable increases in earnings. We show this for England in Figure 1. It appears that housing affordability has worsened considerably since 2001.

One apparent factor contributing to diminished housing affordability is inelastic housing supply (Gyourko et al., 2013). The inability to extend the housing stock stems from both exogenous factors like geography and history, and endogenous factors, such as land use regulations. Research by Saiz (2010) indicates that cities facing inelastic housing supply are constrained by geographical factors, particularly severe land limitations and steep-sloped terrain. While these geographical supply restrictions may explain house price growth in cities like Hong Kong or San Francisco, they are less likely to play a key role in cities with fewer geographical constraints, such as London and Paris. Hence, endogenous supply constraints, such as land use regulations, are critical.

Numerous studies have established a positive correlation between the intensity of land use

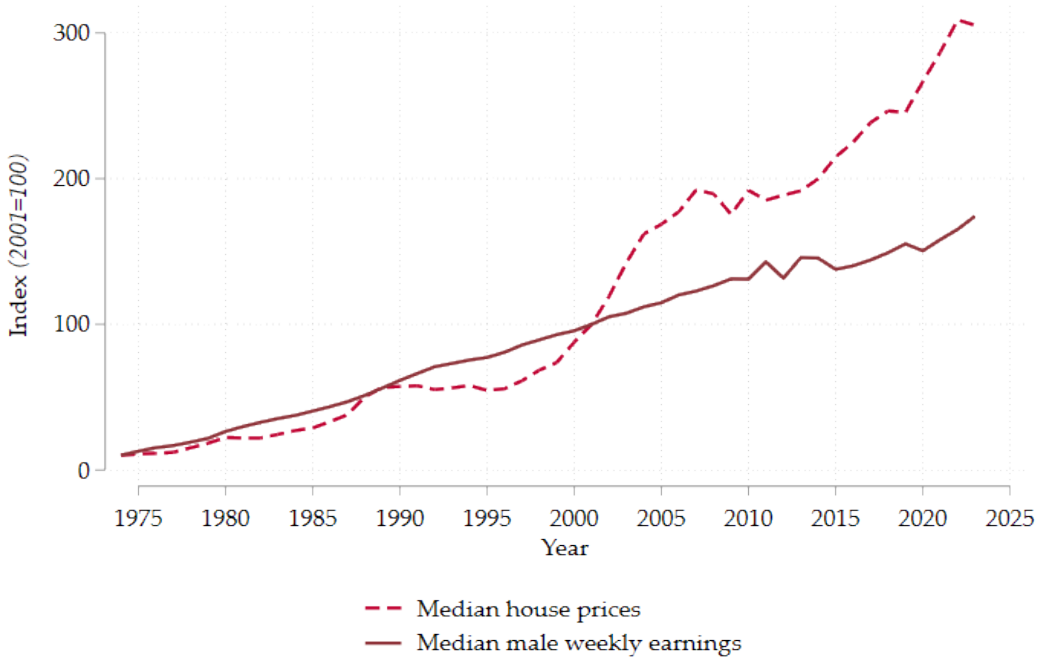


Figure 1 – House prices and earnings over time

Notes: We plot here the real house price and real earnings for England, where we take 2001 as the base year.

regulations and the growth in house prices (Glaeser et al., 2005b; Quigley et al., 2005; Ihlanfeldt, 2007; Glaeser and Ward, 2009; Hilber and Vermeulen, 2016; Severen and Plantinga, 2018). However, it is important to note that the stringency of land use regulations is likely influenced by the current level of development, as building in dense areas tends to be more costly (see Hilber and Vermeulen, 2016). Additionally, the composition of the local population may also contribute to local restrictiveness. The 'homevoter hypothesis', introduced by Fischel (2001), posits that homeowners use their voting power to influence local land-use policies in ways that restrict development. This is particularly the case when homeowners make up a high share of the population in a given area. By supporting more restrictive policies, they aim to preserve or increase their property values by protecting their own views, preventing declines in local amenities (such as the loss of open space to new development, see e.g. Turner, 2005) and by limiting the supply of housing. In doing so, homeowners often prioritise their personal interests over broader economic or social considerations.

Although the homevoter hypothesis is intuitive, the evidence in support of it is somewhat underwhelming. The first aim of this paper is, therefore, to provide more support for the hypothesis that the local tenure composition may affect the intensity of land use regulations and, in turn, new housing construction and house prices. We collect data from England on tenure shares from the decadal censuses since 1981 at the Local Authority (LA) level, of which there were 354 in 2009 in England. We combine this with data on planning applications going back to 1979. Following the literature (see Hilber and Vermeulen, 2016; Cheshire et al., 2018; Hilber and Mense, 2025), we use the number of refused planning applications of major residential projects as our main indicator of regulatory restrictiveness, but we show robustness to other measures. Overall, this provides us with 40 years (1981-2021) of data.

Land use regulation is highly endogenous to the homeownership rate within a local authority (LA), as areas with a higher share of homeowners often differ systematically from others. These LAs typically have higher average incomes, are more likely to be located in suburban areas, and may therefore follow distinct economic or demographic trends. As a result, differences in land use regulation may partly reflect underlying characteristics correlated with homeownership, rather than causal effects. Finding valid instruments for homeownership is a formidable challenge that, to the best of our knowledge, remains unresolved in the existing literature.

We improve on the existing literature by exploiting the Right-to-Buy (RtB) policy implemented in 1980. Under the RtB scheme, eligible tenants have the opportunity to buy their council property at a (strongly) discounted price. The discount is determined based on factors such as the length of the tenancy, the type of property, and its market value. From the start of

the scheme in 1980, more than 3 million housing units owned by local councils have been sold, implying a reduction of the council housing stock of over 60% (Murie, 2022). We then construct an instrument using the initial share of council housing in 1971 (pre-policy) and using the total number of RtB-sales in each year as the shifter in local homeownership. We are well aware that this instrument is subject to the critique that the initial share of council housing in 1971 might be endogenous and correlated to unobserved trends (Goldsmith-Pinkham et al., 2020; Borusyak and Jaravel, 2023).

In all specifications, we include local authority fixed effects and additionally control for the time-varying influence of the 1971 share of council housing. Our identification strategy thus relies on non-linear changes in the popularity of the RtB policy over time. We demonstrate that aggregate RtB sales exhibit substantial year-to-year variation, driven by unforeseen fluctuations in maximum cash grants and other discount-related rules affecting the price RtB households pay relative to market value. Given that this strategy may raise concerns, we conduct an extensive set of robustness checks and also implement an alternative identification strategy following Oster (2019). We highlight two key robustness checks here. First, we show that accounting for non-linear effects of the initial 1971 council housing share has no substantive impact on the results. Second, including controls for the local housing demand—proxied by average earnings within each local authority—does not materially change the estimates.

Applying this strategy, we find that a standard deviation increase in homeownership increases the share of refused planning applications by about 0.7 standard deviations, which is a non-negligible effect. We also find that increased homeownership rates have a dampening effect on housing construction, with annual dwelling growth approximately 0.6% lower for a one standard deviation increase in homeownership. This effect may appear modest, but it accumulates over time. The impact on land development is more pronounced: the share of developed land declines by about 0.25 standard deviations for each standard deviation increase in local homeownership. This suggests that while homeowners may not be able to block new housing entirely—possibly due to national construction targets such as those listed in the National Planning Policy Framework—they are effective at steering development away from their immediate surroundings. This interpretation is reinforced by evidence showing that higher homeownership shares are strongly associated with rising house prices, whereas the effect on rents is both smaller and statistically insignificant.

We hypothesise that owner-occupiers influence local planning decisions through two main channels. First, residents can exercise their influence by voting. Elected councillors have the authority to make planning decisions, and based on previous research, it appears that the Conservative

party is less permissive in allowing for new developments (see e.g., Cheshire et al., 2018). Indeed, we find that homeownership increases local votes for the Conservative party. Second, concerned homeowners can try to influence planning decisions directly. This could involve meeting with councillors as well as submitting objections to planning applications. This will inevitably delay the planning process, which is indeed supported by the data.

We then move to developing a dynamic spatial equilibrium model. We incorporate development decisions and a durable housing stock into a spatial equilibrium model with homeowners and renters. A higher homeownership share curtails new development and raises local house prices. Simultaneously, regulation may enhance local amenity levels, which is differently valued by homeowners.

In a counterfactual analysis, we demonstrate that when setting the homeownership rate to zero, this leads to more permissive land use regulation and increases overall welfare by $X\%$. At the same time, it decreases welfare inequality between renters and owners by $X\%$. The exacerbation of negative welfare effects of land use regulations through increased homeownership levels arises because in the most attractive locations with the highest demand regulations will be the most stringent.

Related literature. Our paper relates to several strands of literature. First, it contributes to a small literature studying the impact of homeownership on regulation. Hilber and Robert-Nicoud (2013) provide cross-sectional evidence that increased homeownership levels—as well as higher levels of development—are associated with increased restrictiveness in the US, although the estimated effect of homeownership is not statistically significant. Dehring et al. (2008) finds that higher shares of homeowners are associated with higher vote shares in a referendum for a public stadium in Texas, which arguably increases property values. Jedwab et al. (2022) study skyscrapers across the world and use the difference between the actual and predicted number of skyscrapers based on the country's GDP as a measure of regulatory restrictiveness. They do not find support for the homevoter hypothesis: cities with higher homeownership rates do not have fewer tall buildings.

Second, our paper ties in with a broader literature investigating the effects of regulation on cities. Most prior research examining the impacts of land-use regulation has centred on the effects of supply constraints. These studies suggest that such restrictions are linked to increasing housing costs, a significant decline in new construction activity, and rapid price appreciation (Mayer and Somerville, 2000; Glaeser et al., 2005a,b; Green et al., 2005; Ihlanfeldt, 2007; Hilber and Vermeulen, 2016; Severen and Plantinga, 2018; Hilber and Mense, 2025). The effect of

supply constraints on house prices is particularly pronounced in English cities, where land-use regulations tend to be highly restrictive (Hilber and Vermeulen, 2016). Other evidence from England, provided by Cheshire et al. (2018), indicates that land-use restrictions may result in increased vacancy rates and longer commute times. In Boston, Glaeser and Gottlieb (2009) find that local constraints within the city do not increase land prices due to the availability of close substitutes, yet density levels are deemed insufficient from a welfare point of view. Koster et al. (2012) suggest that the costs of regulation for homeowners or developers, known as 'own-lot effects', can be substantial, amounting to up to 10% of the housing value. Turner et al. (2014) evaluate the own-lot and amenity effects of land-use regulation in the U.S., finding that while own-lot effects are significant, there is no evidence supporting the presence of amenity effects related to regulation, implying negative welfare consequences of land-use regulation. Harari (2020) demonstrates that the layout of cities is crucial in India. Less compact urban developments lead to longer intra-city travel distances and a lower quality of life. As land-use regulation in India tends to encourage less compact urban forms, it diminishes urban accessibility and welfare. In the context of our paper, we emphasise that in all these studies, land use regulations are a given, rather than being at least partly a result of endogenous decisions by homeowners.

Third, we contribute to an emerging literature on (dynamic) spatial equilibrium models measuring the effects of land use constraints. Hsieh and Moretti (2019) develop a parsimonious spatial equilibrium model to examine the overall impacts of labour misallocation stemming from strict land use regulations. Their analysis reveals that housing supply constraints resulted in a 36% decline in aggregate US growth between 1964 and 2009. These substantial effects are likely attributed to the fact that the most productive cities, such as New York and San Francisco, also exhibit the most stringent land use planning regulations. Dericks and Koster (2021) also document large productivity effects of land use regulations. Using the Blitz as exogenous variation in density restrictions and employing a spatial quantitative equilibrium model, they demonstrate that density constraints exert a significant downward effect on London's GDP by diminishing agglomeration economies. Duranton and Puga (2023) develop a model where human capital spillovers drive entrepreneurship and learning, influencing both land use regulation and welfare outcomes. Incumbent residents restrict city expansion through planning regulations to balance urban costs against productivity gains. The study shows that these regulations, particularly in highly productive cities, limit in-migration and raise housing costs. Relaxing land use restrictions, especially in constrained metropolitan areas, could significantly boost aggregate output and consumption. The above papers typically take an omnibus measure of land use constraints. By contrast, Koster (2024) focuses on a specific type of land use regulation, namely greenbelts.

He develops a spatial equilibrium model including amenities, housing supply, traffic congestion externalities, agglomeration forces, productivity, and household location preferences. The implementation of the Greenbelt policy results in positive amenity effects but significantly diminishes housing supply. The findings indicate that greenbelts enhance overall welfare due to the substantial amenity effects. However, they also exacerbate housing affordability issues by restricting housing supply, which is in line with the reduced-form literature on the effect of housing supply constraints on house prices. Overall, the papers do not address the issue that land use regulations may be impacted by differences in housing tenure across space.

Last, our paper adds to a vast literature on the effects of homeownership on local communities. Hoff and Sen (2005) demonstrate that homeownership amplifies incentives for civic engagement (such as endeavours to combat crime, enhance local governance, and so forth). As a result of within-community externalities, cities become segregated: affluent individuals reside in thriving homeowner communities, while economically disadvantaged individuals inhabit dysfunctional renter communities. Other papers also provide micro-foundations for why increased homeownership can improve local neighbourhoods (DiPasquale and Glaeser, 1999; Field, 2007; Arbel et al., 2017; Sodini et al., 2023). In line with these findings, Hausman et al. (2022) document positive externalities of homeownership on prices of nearby properties using exogenous variation in homeownership in Israel. Our paper adds to this literature by arguing that land use regulations can lead to costs related to NIMBYism and exacerbate spatial segregation among local communities, but at the same time increased homeownership shares may enhance local amenity levels.

The remainder of this paper is structured as follows. Section 2 outlines the institutional setting and data used in the paper. Section 3 documents the reduced-form evidence on the effect of homeownership on the local permissiveness of land use regulations. Section 4 presents our dynamic spatial equilibrium framework. Section 5 concludes.

2 Data and context

2.1 Local planning in England

The English land use planning system stands out internationally due to its discretionary, case-by-case structure known as "development control". This contrasts with the rule-based zoning and master plan systems employed in most other countries. In those countries, zoning ordinances or legally binding master plans delineate land uses in advance, and as long as a proposed development complies with the prescribed rules, approval is typically granted as a matter of right. In England, by contrast, virtually all forms of development—defined broadly to include

not just new construction but also changes in use or intensification of use—require explicit permission from the Local Planning Authority (LPA).

At the heart of the English planning system is the Town and Country Planning Act 1947, which nationalized development rights and established a presumption against development unless explicit permission is granted by the state. Planning applications are assessed in relation to each LPA's Local Plan and the National Planning Policy Framework (NPPF), but both sets of documents are interpreted flexibly. Decisions must weigh "material considerations", and the process is explicitly discretionary.

The process is deeply embedded in public consultation and political negotiation, making it highly sensitive to local political pressures and stakeholder objections. The planning system offers two distinct mechanisms through which local residents can influence development outcomes. The first is electoral: elected councillors, who are ultimately responsible for local planning decisions, are directly accountable to their constituents. The second is procedural: through statutory public consultations, residents can submit objections, participate in hearings, and influence planning officers or committees more directly. Together, these channels create fertile ground for NIMBY ("Not In My Backyard") dynamics, as locally organised opposition to new development can be translated into political and administrative pressure.

In practice, most planning decisions are taken by local planning committees operating within each LPA. These committees are composed of locally elected councillors and are required to reflect the political composition of the council as a whole. While statutory planning frameworks provide guidance, councillors often rely heavily on professional planning officers—career civil servants employed by the LPA—who assess applications, prepare recommendations, and manage the day-to-day operation of the development control system. This reliance partly reflects the technical complexity of planning law and practice. As a result, planning officers exercise substantial de facto influence over outcomes on the ground, even though decisions are formally taken by elected members. Importantly, however, both councillors and officers remain embedded in a political environment shaped by local electoral incentives and resident opposition to development: councillors are directly exposed to NIMBY pressures through electoral accountability, while planning officers are indirectly exposed through their appointment, oversight, and interaction with elected members.

Taken together, this institutional structure provides a natural channel through which local tenure composition—the share of homeowners versus renters—can shape regulatory outcomes, helping to rationalise why higher homeownership rates are associated with higher refusal rates and longer decision times in the reduced-form evidence below.

An additional institutional feature further enhances the suitability of the English case for a welfare analysis of the impact of homeownership and NIMBYism. Unlike in many other countries, private sector rents in England are largely unregulated, and even social housing rents, though heavily subsidized, tend to move broadly with market rates (Hilber and Mense, 2025). This implies that renters in England are more exposed to housing market signals and amenity changes than in systems with rigid rent controls. Consequently, English renters may behave differently from their counterparts in more regulated settings, making the contrast between renters' and homeowners' preferences more empirically discernible.

The development control system is complicated by an array of additional constraints. These include height limits, protected view corridors, conservation area designations, listed buildings (a more restrictive form of preservation), and expansive Green Belt zones that encircle many urban areas (Koster, 2024). Originally designed to prevent urban sprawl, Green Belts now severely limit outward growth, particularly in the high-demand areas around London and the South East. Simultaneously, upward development is often blocked by height restrictions and preservation policies. Together, these layered constraints significantly reduce the flexibility of land markets and restrict both horizontal and vertical expansion. Planning approvals also depend on complex negotiations between developers and LPAs through mechanisms such as Section 106 agreements, which mandate developer contributions to infrastructure or affordable housing. These negotiated obligations, combined with regulatory uncertainty, increase development costs and delay timelines—further contributing to England's well-documented housing supply inelasticity and worsening affordability challenges (Hilber and Vermeulen, 2016; Drayton et al., 2025; Hilber and Mense, 2025).

2.2 Data

Our data are constructed for England's Census years 1981, 1991, 2001, 2011, and 2021, and are derived from several sources. We focus on Local Authorities, of which there are 354 in England.[1]

First, our measure for regulatory restrictiveness is from the Department for Levelling Up, Housing and Communities's (DLHC) Planning Statistics. Consistent with the existing literature (see e.g., Hilber and Vermeulen, 2016; Cheshire et al., 2018), our primary metric is the share of refused planning applications for major residential projects, obtainable for each Local Authority (LA) annually. Major residential developments comprise 10 or more dwellings that an LA received during the 'development control' process in any given year. We also utilise data

[1]We adopt the pre-2009 definition of Local Authorities (LAs). Subsequent to 2009, the number of LAs has been reduced to 326. Where necessary, we adjust post-2009 values to align with the LAs in 2009, considering weights based on the 2009 population or number of dwellings. Additionally, we compile an annual panel dataset using supplementary sources on housing construction and sales under the Right-to-buy scheme. For further information, please refer to Appendix B.1.

on the proportion of 'delayed' major dwelling planning applications, defined as the percentage of applications that exceed a duration of 8 weeks before a decision is taken. Additionally, for some sensitivity analyses, we also collect information on decisions related to minor residential and commercial developments.

Second, the local homeownership rate in each LA is derived from Population Censuses. Also, from the Censuses we derive the share of council housing, housing owned by housing associations, and the share of private-rental housing. From the Censuses, we also get the number of dwellings in each LA.

Third, to construct the instrument for homeownership, we use data from the DLHC on the total number of right-to-buy sales in England since the inception of the policy in 1980. Based on this we count the aggregate growth (decline) in owner-occupied (council) housing units, abstracting from local new construction and demolitions. In other words, we assume that right-to-buy sales result in a direct conversion of council housing units into owner-occupied housing units. The details and validity of this instrument will be thoroughly examined later on. Based on the initial number of owner-occupied housing units in 1971, we then predict the ratio of owner-occupied to council housing units in each LA for each year.

Fourth, we obtain information on the share of developed land from the Land Cover maps available for the years 1990, 2000, 2007, 2015, 2019, 2020, and 2021. To mitigate coding errors, we calculate the share of developed land within each postcode, which is the smallest available geographical unit. For each consecutive land cover year, we impose a non-decreasing constraint on the share of developed land per postcode. For the intermediate years, we interpolate the postcode-level shares and subsequently aggregate them to the level of local authorities. Alternatively, we obtain the share of urban land from Li et al. (2021), which is determined at a 1 km spatial resolution by combining historical data, satellite observations, and urban dynamic modelling using a Logistic-Trend-CA model. Urban land extent is reconstructed from 1870 to 1991 using hindcasting, and observed from 1992 to 2013 via nighttime light satellite data.

Fifth, we collect data on property values. House prices are based on transaction data from the Council of Mortgage Lenders for the years 1974-1995 and from the Land Registry for 1995–2021. We calculate average prices at the local authority level. Rent data are sourced from the Survey of English Housing (SEH) for 1993–2006 and from the Bricks & Logic database for 2007-2021. To estimate rents prior to 1993, we use reported moving-in dates and adjust historical rents by the retail price index. Because rental data are sparse in the early years and for some LA-year cells, we impute rents when there are fewer than five observations using a random forest model. This model draws on information about private-registered provider rents, social sector rents, house

prices, year, and geographic location (latitude and longitude), and performs exceptionally well as the correlation between predicted and observed rents in the training sample is 0.997.

Sixth, we use data on local earnings, which we base on male weekly earnings spanning from 1974 to 2021. Earnings data between 2004-2021 are sourced from the Annual Survey of Hours and Earnings (ASHE), while data for 1974-2004 are obtained from the New Earnings Survey (NES). While ASHE data is available at the LA-level, NES data for earlier years is only accessible at the county and London borough levels. To ensure consistency, we geographically match all earnings data to the LA-level and adjust nominal earnings figures using the Retail Price Index to derive real earnings. For a more comprehensive understanding of the earnings data and methodologies employed, we refer to Hilber and Vermeulen (2016). We also construct a labour demand shock, based on the employment in 7 industries in 1971 and the development in each of these industries between 1971 and 2021.

Seventh, we rely on data related to the political composition of LAs and local vote shares. We collected local election data from various sources: (i) the British Local Election Database (1889-2003) compiled by Rallings and Thrasher (2004), (ii) the Local Election Handbooks (1999 to 2021), (iii) the Local Elections Archive Project (LEAP) (2006 to 2010), and (iv) the BBC (2009 to 2011).

Finally, we acquire weather variables, including minimum and maximum temperatures, hours of sunshine, and rainfall (in mm), from the Met Office for each LA in a specific year. Additionally, for each LA we collect data on the age composition, the share of people with (permanent) disabilities, and the share of people with an education degree from the Censuses.

2.3 Descriptives statistics

Given 5 waves of data and 354 LAs, we have 1,770 observations.[2] We show descriptive statistics in Table 1. The mean number of planning applications across Local Authorities (LAs) is approximately 20.[3] Out of 88 LA-year observations (5%), no major dwelling planning applications were recorded, leading to the exclusion of these observations from the regressions. The average refusal rate across LAs stands at about 25%. The average refusal rate across England follows a highly cyclical pattern with higher refusal rates during periods of increased housing demand (see Figure B1 in Appendix B.2). The rate of planning applications facing delayed decisions is rather high, standing at 68%, and exhibits less cyclicality compared to the share of refused planning

[2]For the LA Redcar and Cleveland, we do not observe the share of Labour seats in 1981 and 1991, reducing the number of observations for some specifications to 1,768.

[3]According to construction statistics, there have been about 6,839,430 million dwellings constructed between 1979-2021, while the total number of granted major dwelling applications has been 239,782. If we assume that two-thirds of the properties are constructed in major projects, the average size of a project is about 20 dwellings.

Table 1 – Descriptive statistics

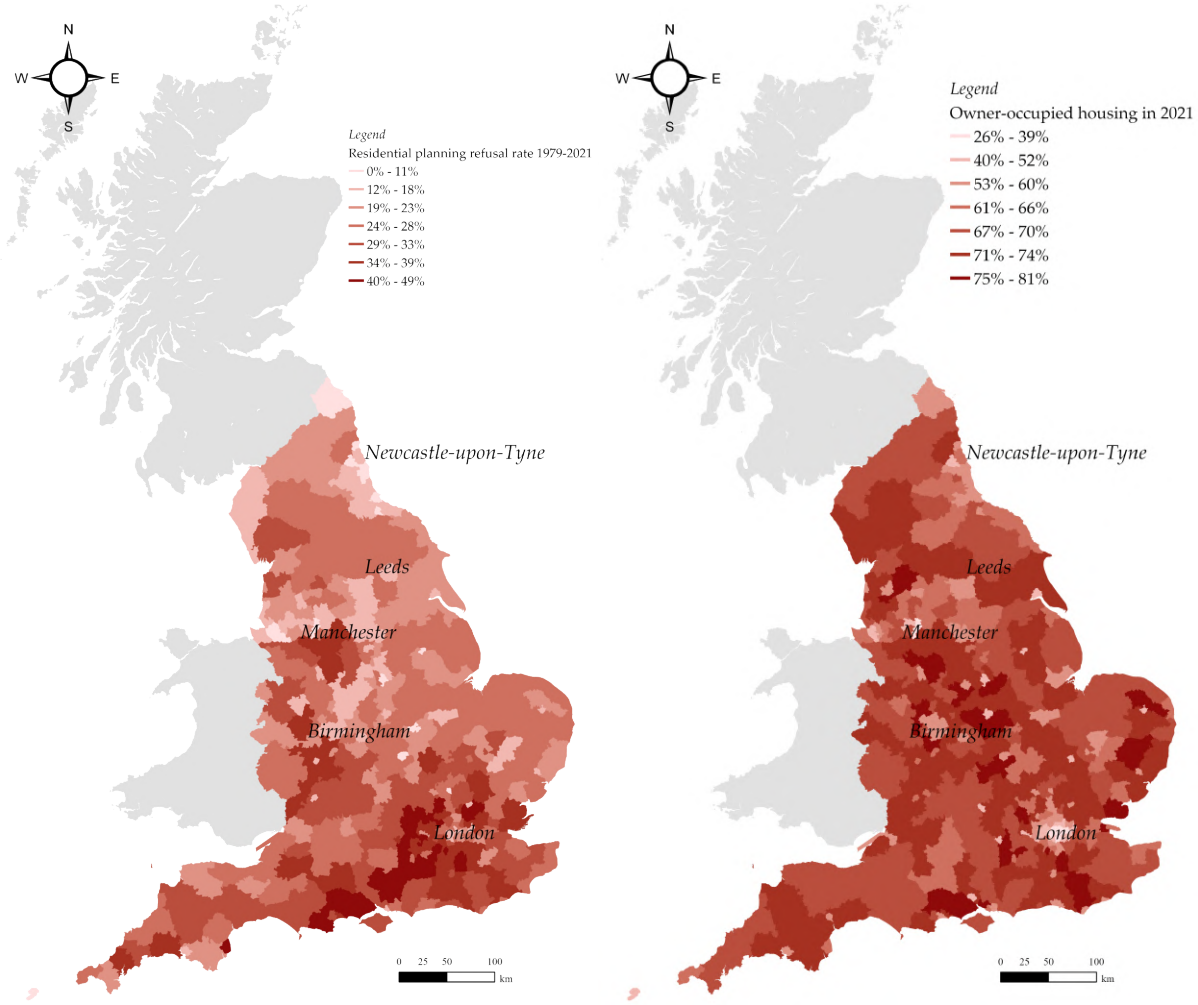
	Mean	Standard deviation	Min	Max
	(1)	(2)	(3)	(4)
Residential planning refusal rate (%)	24.33	17.42	0	100
Residential planning delay rate (%)	67.90	20.11	0	100
Residential planning applications	19.30	14.66	1	130
Share owner-occupied housing (%)	66.86	10.78	16.54	89.52
Share council housing (%)	13.83	10.47	0.410	74.86
Share housing association units (%)	5.429	4.247	0.0849	23.06
Share private rent (%)	13.88	6.355	1.309	43.59
Number of dwellings	56,470	38,027	942	423,529
Mean house price (£)	160,113	158,306	15,183	2.006e+06
Mean annual rent (£)	8,109	6,060	667.7	88,822
Share of Conservative seats (%)	41.69	24.05	0	100
Share of Labour seats (%)	31.70	27.92	0	100
Share of liberal-democrat seats (%)	13.96	16.17	0	92.31
Share of other seats (%)	12.65	17.01	0	100
Male weekly earnings (£)	468.1	238.4	89.54	1,793
Predicted employment	74,232	57,811	1,742	605,258
Predicted share of owner-occupied housing (%)	76.49	13.10	6.295	95.05

Notes: The number of local authorities included is 354. Given the 5 waves of data we have, this culminates to 1,770 observations. For 88 observations the number of major dwelling planning applications is zero so these observations will be dropped in the baseline regressions. The residential planning delay rate is defined as the share of planning decisions that took more than 8 weeks. We adopt the pre-2009 definition of Local Authorities (LAs). Subsequent to 2009, the number of LAs has been reduced to 326. Where necessary, we adjust post-2009 values to align with the LAs in 2009, considering weights based on the 2009 population. Here, we do not report the weather variables, including minimum temperature, maximum temperature, sunshine hours, and rainfall. We also do not report demographic variables, encompassing the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with an education degree. Those values are available upon request.

applications. Therefore, this suggests that LAs are not necessarily overwhelmed during periods of high demand, as there is no significant increase in the delay rate.

Going back to Table 1, it appears that England has a substantial share of owner-occupied housing, amounting to 67%. We note that this figure is unweighted; when weighted by the number of dwellings in each LA, the share of owner-occupied housing decreases slightly to 63%. Council housing accounts for 14% of the total dwelling stock, while housing units owned by housing associations comprise just over 5%. Interestingly, the share of owner-occupied housing was already increasing before the RtB policy from about 40% in 1961. After the implementation of the policy, owner-occupied housing kept increasing until it reached its peak in 2000 at 69%. As expected, the share of council units has steadily decreased since 1980, while England's private rental sector is relatively small at 14%.

The Conservative Party holds the largest share of local council seats, with over 40% representation across LAs. There is considerable variation in weather conditions across England in terms of temperatures, rainfall, and hours of sunshine. The average weekly earnings for males are



(a) Residential planning refusal rate (b) Share owner-occupied housing

Figure 2 – Spatial variation

Notes: These are maps based on local-authority level data. The residential planning refusal rate is based on data spanning 1979-2021. The share of owner-occupied housing is from 2021.

approximately £470. Given the likelihood of endogeneity in earnings, we will also employ an alternative proxy for local demand based on predicted employment. This entails a classical shift-share instrument derived from local employment shares in 7 sectors in 1971 and the subsequent aggregate employment growth in these sectors in England.

The predicted share of owner-occupied housing, which serves as the instrument to address endogeneity concerns regarding the share of homeownership, averages 77%. It is worth noting that this figure appears higher because it is the share with respect to council housing units only.

We provide supplementary descriptive statistics in Appendix B.2, highlighting the significance of electoral cycles, along with histograms depicting the key variables.

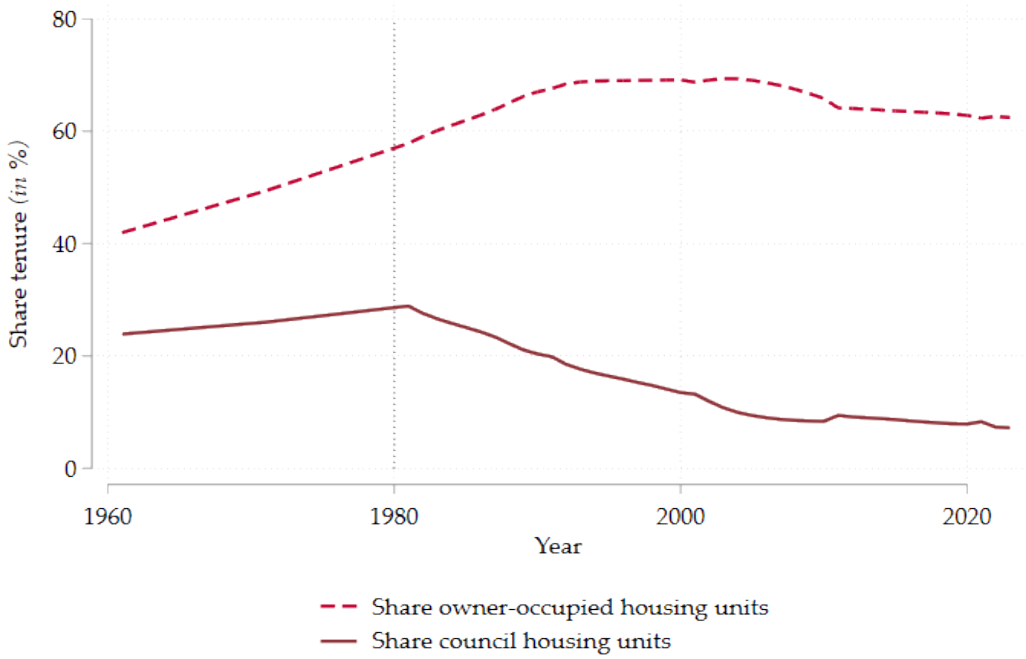


Figure 3 – Housing tenure over time

Notes: We plot here the share of owner-occupied housing in England. For the non-census years we interpolate the data using additional statistics from the DLHC, which provide dwelling and tenure statistics since 2001. Before that we use data on right-to-buy sales and construction by tenure type. The vertical dashed line in 1980 indicates the start of the Right-to-Buy scheme.

3 Reduced-form evidence

3.1 Methodology

Baseline estimation. Our main focus here is to demonstrate the impact of homeownership on regulatory restrictiveness, as proxied by the residential planning refusal rate. We will also show that homeownership affects new construction, conversion to developed land, and property values. Additionally, we seek to illustrate that this effect occurs through two channels: homeowners voting for parties that endeavour less permissive planning regulation, and homeowners directly influencing the planning process, which inevitably leads to delays in planning application decisions.

Let r_{it} be the planning refusal rate in LA i in year t and h_{it} is the share of homeownership in an LA i in year t. We then estimate:

$$r_{it} = \beta_1 h_{it} + \beta_2 x_{it} + \mu_i + \mu_t + \epsilon_{it} \tag{1}$$

where β_1 represents the key coefficient of interest, x_{it} denotes a collection of time-varying LA-level variables, such as weather variables, and local housing demand. Additionally, μ_i and μ_t capture LA and year fixed effects, respectively.

This equation is subject to a number of concerns. First, one may be concerned about reverse causality because increased regulatory restrictiveness may increase prices, which in turn may affect the share of owner-occupied housing. Second, the regression above likely suffers from omitted variable bias. We may expect a negative relationship between homeownership and local housing demand, as developers tend to invest more in areas with higher demand pressure. These new developments are more frequently rental projects.[4] Due to the cyclical nature of refusal rates (see Figure B1 in Appendix B.2), which are higher in areas and times with greater demand, inadequately controlling for local demand likely leads to an underestimation of the impact of homeownership on regulation.

Right-to-Buy as an instrument for homeownership. We therefore propose to instrument for h_{it}. We exploit exogenous variation in local homeownership rates due to the Right-to-Buy (RtB) policy, which was introduced during the tenure of Margaret Thatcher's Conservative government in 1979. It afforded tenants of public housing the opportunity to buy their homes at a substantial discount relative to market prices (ranging from 30% to 70%, see Koster and Pinchbeck, 2024). Under RtB, eligible tenants, irrespective of income, benefit status, or other circumstances, must have resided in public housing for at 3 years to be eligible. Households must submit an application to their LA, which then assesses the home's market value and determines the applicable discount. Applicants have the option to contest the LA's valuation, leading to a secondary assessment by a District Valuer. Upon approval, households opting for RtB must secure a mortgage and cover associated legal fees, survey expenses, and possibly Stamp Duty Land Tax. Upon completion of the purchase, beneficiaries become ineligible for Housing Benefit and typically forfeit their ability to reapply for public housing. Today, RtB largely persists in its original form for tenants of publicly-owned housing although RtB discounts have been reduced over time. As of now, approximately 3 million households have exercised their right to buy (Murie, 2022).

By utilising data from the DLHC, we gather the cumulative count of Right-to-Buy (RtB) sales in England since the policy's inception. Subsequently, we anticipate the most pronounced shifts in homeownership within LAs that initially exhibited a high proportion of council housing. Let us define:

$$\begin{aligned} \hat{n}^o_{it} &= n^o_{i1971} \times \left(\frac{n^o_{1971} + \text{rtb}_t}{n^o_{1971}} \right), \\ \hat{n}^c_{it} &= n^c_{i1971} \times \left(\frac{n^c_{1971} - \text{rtb}_t}{n^c_{1971}} \right), \end{aligned} \tag{2}$$

[4] In our data, the cross-sectional correlation between the share of owner-occupied housing and a Bartik demand shock is -0.405. Conditional on LA and year fixed effects, the (partial) correlation is -0.281.

where n^o_{i1971} and n^c_{i1971} are, respectively, the number of owner-occupied and council housing units in 1971 in i, n^o_{1971} are the total number of owner-occupied dwellings in 1971 in England, and rtb_t are the cumulative number of RtB sales since the policy's inception in 1980. Then, our instrumental variable is defined as:

$$z_{it} = \log\left(\frac{\hat{n}^o_{it}}{\hat{n}^o_{it} + \hat{n}^c_{it}}\right), \tag{3}$$

so z_{it} represents the log of the predicted share of owner-occupied housing vis-à-vis council housing as a result of an increasing number of RtB sales in England. In Figure B4 in Appendix B.2 we show that, on average, the predicted share of owner-occupied housing increases from about 65% in 1980 to about 80% in 2021.

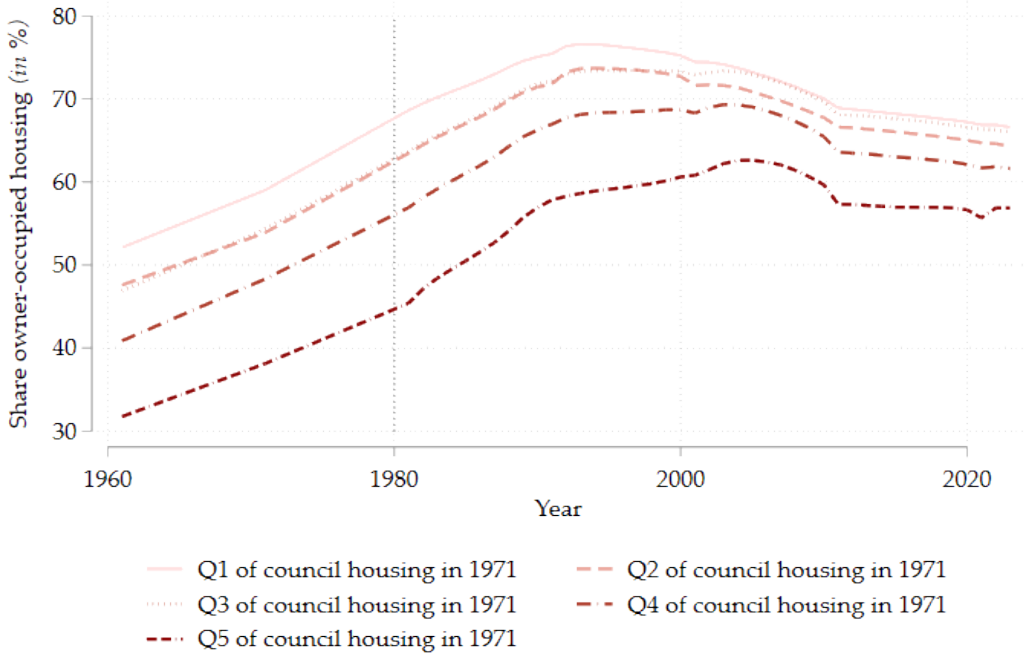
The main concern with the instrument is that local authorities with high initial shares of council housing may face lower housing demand for unrelated reasons, prompting planning authorities to approve nearly all construction projects in subsequent years. This suggests that the initial shares are correlated with changes in refusal rates. This relates to a broader criticism that the identification may at least partly rely on initial housing tenure shares, which could be endogenous. (Goldsmith-Pinkham et al., 2020; Borusyak and Jaravel, 2023).[5] One thing one can do is to inspect pre-trends in initial shares. We show in Figure 4 supporting the exogeneity of initial shares of council housing in 1971 by plotting trends in homeownership and earnings from respectively 1961 and 1974 onwards (so before RtB was implemented) by quantile of the share of council housing in 1971. We find strong support that LAs with different shares of council housing exhibit parallel pre-trends in homeownership and earnings, which only start to differ years after the RtB was implemented.

One might still question whether the absence of pre-trends, as suggested by the previous evidence, is sufficiently convincing. To address this concern, we extend equation (1) as follows:

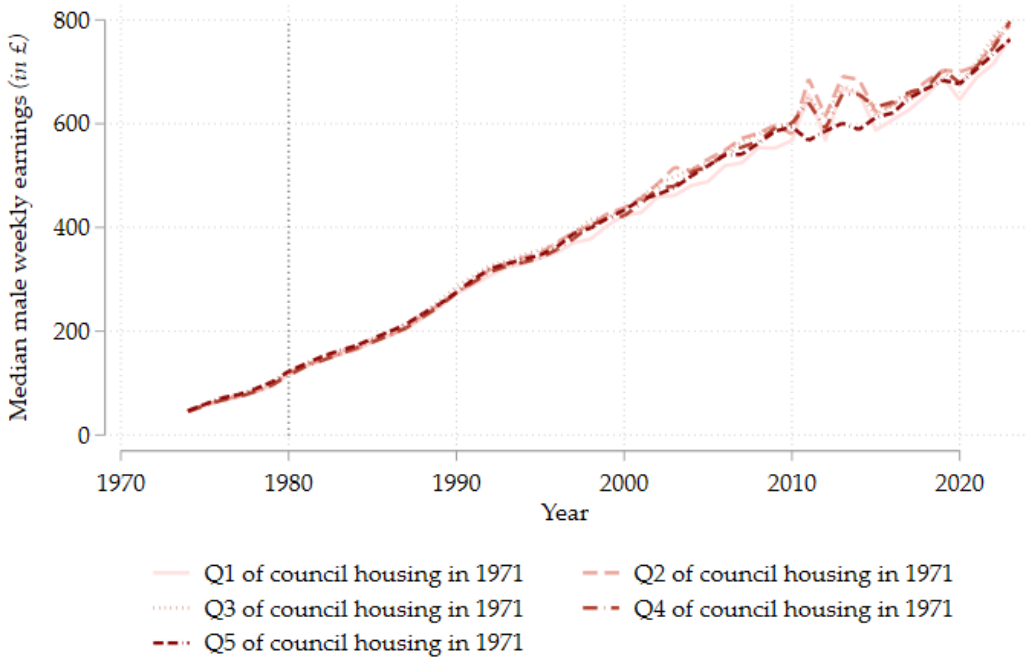
$$r_{it} = \beta_1 h_{it} + \beta_2 x_{it} + \beta_{3t} c_{i1971} + \mu_i + \mu_t + \epsilon_{it} \tag{4}$$

where c_{i1971} denotes the share of council housing in 1971 and β_{3t} captures its time-varying effect on the refusal rate. Further, we include controls for a predicted demand shock, derived from industry shares in 1971 and aggregate industry growth thereafter (see Hilber and Vermeulen, 2016, for details). Essentially, the demand shock accounts for shifts in local demand resulting

[5]There might be particular concerns about council housing being concentrated in low-demand areas in 1971. However, the correlation between log house prices and the share of council housing in 1971 is only −0.075. Indeed, a closer examination of Figure B5a in Appendix B.2 reveals considerable seemingly random variation in the council housing distribution across LAs.



(a) Owner-occupied housing



(b) Earnings

Figure 4 – Trends by quantile of council housing in 1971

Notes: These are based on local-authority level data for annual data from 1961, 1971, and 1981-2001 for homeownership and 1971-2021 for earnings.

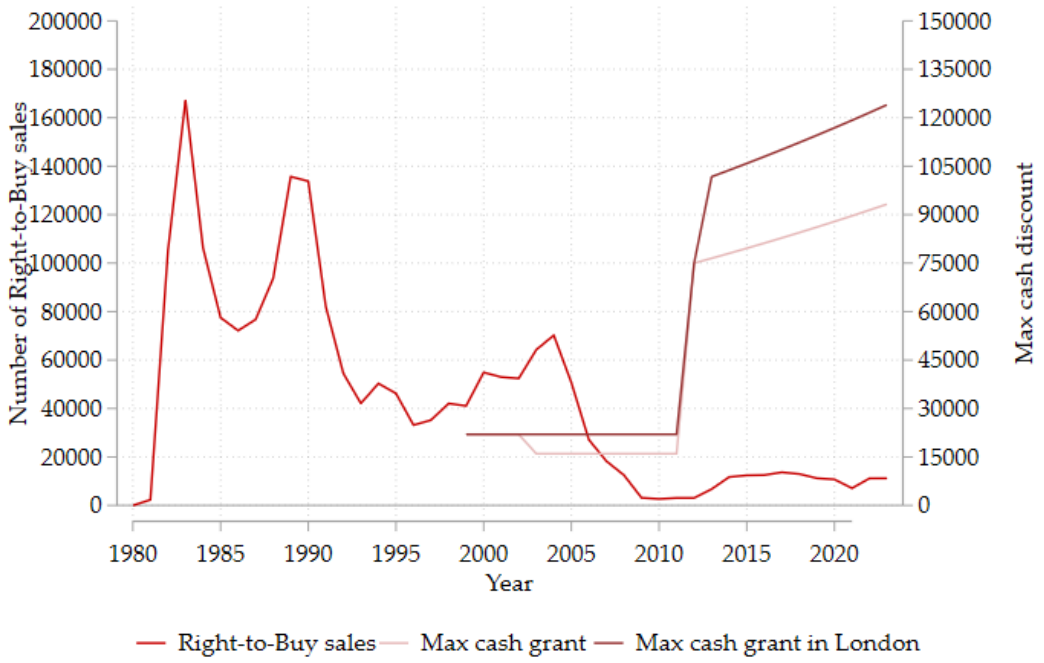


Figure 5 – Right-to-Buy sales over time

Notes: We plot the number of RtB sales per year (left y-axis) and the maximum cash grant (right y-axis). Before 1999, there was no maximum cash grant, but maximum discounts of 50% for houses and 70% for flats were in place throughout the study period.

from aggregate changes in the industry structure in England.

Controlling for the time-varying impact of c_{i1971} implies that we identify the effect of non-linear changes in aggregate RtB sales that locally changed the composition by housing tenure. Figure 5 shows that the number of RtB sales evolved in a highly non-linear manner over time, with the strongest effects on homeownership occurring in the early years of the sample. The aggregate number of RtB sales appears to respond only gradually to the introduction of maximum cash grants in 1999, which were capped at just £22,000 between 1999 and 2011. RtB activity is also closely linked to house price dynamics: during market downturns, participation in the scheme tends to decline.

In addition to a further set of robustness checks supporting the plausibility of our identifying assumptions, we conduct regression analyses using annual panel data, focusing on the years surrounding the introduction of the RtB policy. Finally, we consider an alternative identification strategy using Oster's (2019) methodology to adjust the estimates for omitted variable bias. Our findings demonstrate that the primary results remain robust.

Effects on housing construction, developed land and property values. Subsequently, we examine the effects of homeownership on construction and house price changes. Hilber and Vermeulen

(2016) demonstrate that regulation prompts higher house prices in more tightly regulated LAs due to a lower supply of housing. Consequently, we expect that, as having more homeowners drives greater regulatory restrictiveness, this will ultimately lead to reduced housing construction and higher home price appreciation. We then move on to estimate:

$$\left\{ \log \frac{n_{it+2}/n_{it}}{2}, d_{it}, \log P_{it}, \log p_{it} \right\} = \beta_1 h_{it} + \beta_2 x_{it} + \beta_{3t} c_{i1971} + \mu_i + \mu_t + \epsilon_{it}, \tag{5}$$

where n_{it} is the number of dwellings in t, d_{it} is the share of developed land, P_{it} is the house price, and p_{it} denotes rents. Once more, we instrument for h_{it} with z_{it}.

Two points merit attention. First, one might worry that the effect of homeownership on restrictions to new housing supply takes time to materialise. To address this, we estimate our regressions using the net change in the number of dwellings relative to $t+2$, and also consider the change relative to $t+10$ as an alternative specification.

Second, regarding the effects on house prices and rents, we acknowledge that the estimated coefficient β_1 in the above specification might also capture favourable amenity effects (such as the preservation of green space) and potentially, positive externalities associated with increased homeownership. We expect the effects on current house prices to be considerably stronger than on rents, given that house prices are inherently forward-looking. In our spatial equilibrium model, we aim to disentangle the supply effects and the amenity effects over time of increased homeownership.

Mechanisms: voting and direct influence. Lastly, we will offer some suggestive evidence on the mechanisms by which homeowners can impact land use regulation. We will explore two mechanisms: voting and homeowners directly influencing planning restrictions. In terms of voting, there is evidence suggesting that England's Conservative Party is less inclined than other parties to permit new construction. The bivariate correlation between the refusal rate and the share of Conservative party seats is 0.26, and previous research has consistently demonstrated a strong positive relationship between the share of Conservative seats and regulatory stringency (see e.g., Cheshire et al., 2018). The second channel is homeowners' direct influence on planning decisions. Although precise proxies for this are hard to define, we posit that direct influence primarily leads to delays in planning decisions, as homeowners may submit petitions and amendments. We then estimate:

$$\{v_{it}, \ell_{it}\} = \beta_1 h_{it} + \beta_2 x_{it} + \mu_i + \mu_t + \epsilon_{it}, \tag{6}$$

Table 2 – Homeownership and regulation

Dependent variable:	Residential planning refusal rate (standardized)				
	Naïve OLS	+Control variables	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)	(4)	(5)
Share owner-occupied housing (sd)	-0.170***	0.012	0.731***	0.752***	0.721***
	(0.056)	(0.048)	(0.275)	(0.284)	(0.277)
Residential planning applications (log)					-0.007
					(0.069)
Male weekly earnings (log)					0.668***
					(0.149)
Share council housing 1971×year dummies			✓	✓	✓
(Share council housing 1971)^2×year dummies				✓	✓
Controls		✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,682	1,682
R^2	0.451	0.461			
Kleibergen-Paap F-statistic			568	581.6	499.5

Notes: (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. In Columns (3)-(5), we instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

where v_{it} is the share of Conservative seats and ℓ_{it} is the share of planning decisions taking at least 8 weeks. We instrument for h_{it} using (3).

3.2 Baseline results

The baseline results are presented in Table 2. The OLS specification yields a negative coefficient for the share of owner-occupied housing. However, as previously discussed, naïve OLS estimates are likely downward biased, since developers may be more likely to submit applications in high-demand areas for larger projects consisting of privately rented properties. When we control for the Bartik demand shock, weather, and certain demographic characteristics, the coefficient becomes statistically insignificant and effectively zero.

In column (3), we implement an instrumental variable approach, using the predicted share of owner-occupied housing based on RtB sales as an instrument. At the same time, we control for interactions between the 1971 share of council housing and year dummies to absorb any time-varying trends related to the initial share of council housing and its influence on residential planning refusal rates. In Appendix C.1, we show that the predicted share of owner-occupied housing has a robust and positive effect on the share of owner-occupied housing. The Kleibergen-

Paap F-statistic indicates that our instrument is sufficiently strong.

Going back to Table 2, Column (3) reveals a strong and statistically significant positive effect of homeownership on regulation: a standard deviation increase in homeownership increases the refusal rate by 0.7 standard deviations. Column (4) we improve on identification by further controlling for the share of council housing in 1971 squared and interact this with year dummies. In this way, we control for non-linear trends in the initial share of council housing that may be correlated to refusal rates and the instrument. The fact that our results are not materially impacted suggest that unobserved trends related to the initial shares in council housing in 1971 is not a major issue.

In the final column of Table 2, we add two potentially endogenous controls. First, one might argue for including the number of planning applications, since local authorities facing a high volume of applications may become overwhelmed and consequently reject more planning proposals. However, we find no significant association between the number of residential planning applications and the refusal rate. In Section 3.3, we further show that homeownership has no discernible effect on the number of planning applications.

Second, we control for the log of average earnings in a local authority. If the Bartik demand shock does not fully capture local demand conditions, earnings may serve as a more direct proxy. While we find that local authorities with higher average earnings are substantially more restrictive, the estimated effect of homeownership on the refusal rate remains unchanged.

3.3 Robustness

Other robustness analyses. Table 3 presents several robustness checks. In column (1), we restrict the sample to areas with council housing shares in 1971 above 28% (the median) to address residual concerns that initial levels of owner-occupied housing might be correlated with underlying regulatory trends. It is reassuring that the point estimate remains largely unchanged, although the standard errors increase. When we further restrict the sample to local authorities with council housing shares above 28% in column (2), we reach a similar conclusion.

Column (3) introduces two other potentially endogenous controls associated with refusal rates: the number of dwellings (reflecting housing construction) and house prices (influencing the cyclicality of refusal rates). Neither variable demonstrates a positive influence on the refusal rate. More importantly, the inclusion of these controls has minimal impact on the baseline estimate, thereby enhancing the credibility of our identification strategy.

In column (4), we test for non-linear effects of the homeownership share. The coefficients imply that the marginal effect is positive for homeownership shares below 75% and then turns slightly

Table 3 – Homeownership and regulation: robustness

Dependent variable:	Residential planning refusal rate (standardized)				
	Above-median council share 1971	Above 75th perc. council share 1971	+House prices, dwellings	Nonlinear effect	Homeownership majority
	(1)	(2)	(3)	(4)	(5)
Share owner-occupied housing (sd)	0.885*	0.738	0.712**	1.636***	
	(0.522)	(0.727)	(0.281)	(0.339)	
(Share owner-occupied housing)^2 (sd)				-9.016***	
				(2.582)	
Majority of homeowners					3.103***
					(0.638)
Dwellings (log)			-0.586		
			(0.628)		
Mean house price (log)			0.121		
			(0.289)		
Share council housing 1971×year dummies	✓	✓	✓	✓	✓
Weather and demographic controls	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓
Number of observations	846	420	1,682	1,682	1,682
Kleibergen-Paap F-statistic	823.2	58.55	206	150.9	20.78

Notes: (sd) denotes variables standardized to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors, shown in parentheses, are adjusted for cross-sectional dependence based on a 10-year time window following Driscoll and Kraay (1998). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

negative; however, it is far from being statistically significantly negative for any homeownership rate observed in the data.

In the preceding analyses, we use the share of owner-occupied housing as the main explanatory variable. This measure is both simple and intuitive, capturing the influence of homeowners in a given area, particularly through channels such as voting and lobbying. One might argue, however, that a more relevant measure would be a dummy variable indicating whether a local authority has a homeowner majority, since a majority may be decisive in blocking new developments. In column (5) we find that a homeownership majority increases the refusal rate by 3 standard deviations.

The majority threshold may be considered restrictive, however, as homeowners tend to exhibit significantly higher voter turnout than renters (see Manturuk et al., 2009; Hall and Yoder, 2022), which amplifies their political influence in local elections and planning decisions—even when they represent less than half of the population. We thus re-estimate equation (4), replacing h_{it} with a dummy variable indicating whether the share of owner-occupied housing exceeds

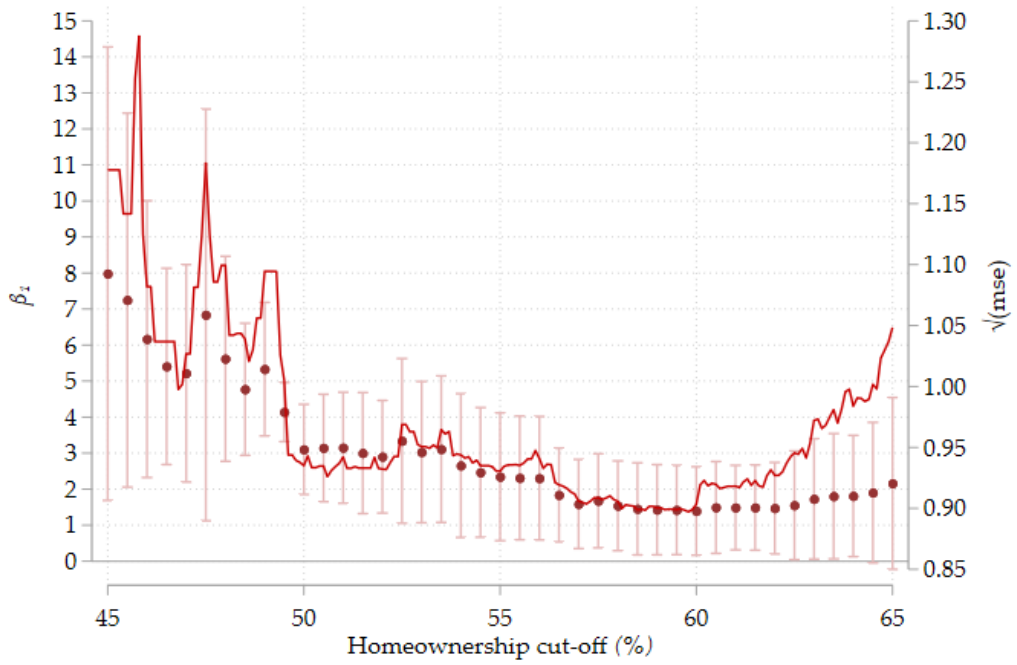


Figure 6 – The homeownership cut-off

Notes: We replicate specification (3) from Table 2 using different homeownership cut-offs. The coefficients are plotted on the left y-axis, with dots representing point estimates and vertical lines denoting the 95% confidence intervals that are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998). The continuous red line represents the RMSE, which is displayed on the right y-axis.

a given threshold. Figure 6 presents the results. The root-mean-squared-error reaches its minimum for homeownership cut-offs between approximately 50% and 65%. However, because the differences in coefficients across thresholds are minimal, we cannot identify a precise tipping point. Nevertheless, the results are robust across this range: a homeownership share of at least 50% is associated with an increase in the refusal rate of roughly 2 to 3 standard deviations.

Annual panel data. In Table 4, we use an annual 42-year panel dataset covering 1979–2021 (see Appendix B.1 for details). This offers a substantial increase in the number of observations and a strong first-stage relationship. The second-stage coefficient in column (1) is slightly smaller than the baseline estimate, although of the same order of magnitude. In column (2), we restrict the sample to the years 1979–1999, the period with the highest volume of RtB sales (recall Figure 5). The resulting coefficient closely aligns with the baseline estimate from the decennial panel.

Column (3) narrows the analysis to the years surrounding the policy's introduction, where most of our identifying variation comes from (see Figure 5). Specifically, we include the years 1979 and 1984. Although we again find a positive effect of homeownership on refusal rates, the coefficient is positive but imprecisely estimated. When we instead use the years 1979 and 1989,

Table 4 – Homeownership and regulation: annual panel data

Dependent variable:	Residential planning refusal rate (standardized)			
	1979-2021	1979 1999	1979 & 1984	1979 & 1989
	(1)	(2)	(3)	(4)
Share owner-occupied housing (sd)	0.335***	0.572*	1.697	0.753*
	(0.119)	(0.313)	(1.448)	(0.428)
Share council housing 1971×year dummies	✓	✓		
Controls	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓
Number of observations	14,573	6,863	616	616
Kleibergen-Paap F-statistic	453.8	342.6	84.20	155.8

Notes: (sd) denotes variables standardized to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. The demographic variables are imputed for the non-census years based on linear interpolation. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are shown in parentheses. They are adjusted for cross-sectional dependence based on a 10-year time window following Driscoll and Kraay (1998) in columns (1) and (2) and are robust standard errors in columns (3) and (4). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

the estimated effect remains positive and closely matches the baseline result reported earlier.

Other dependent variables. In Table 5, we substitute the dependent variable with alternative measures of regulatory restrictiveness. In Column (1), we adjust the dependent variable so that the refusal rate is calculated over applications received between $t-2$ and $t+2$. However, this adjusted measure excludes data from 2021 due to the unavailability of post-2021 planning application data at the time of writing. The point estimate is highly statistically significant and very similar to the preferred specification reported in Table 2.

Column (2) examines the effect of homeownership on the log number of planning applications, finding no effect on the total number of applications. This may go against the initial expectation that higher refusal rates associated with higher homeownership rates may discourage developers from applying in the first place. At the same time, homeowner opposition to large developments may lead to more planning applications, as developers submit multiple or revised proposals to navigate local resistance and planning regulations. Further, homeowners may oppose dense, affordable housing but support high-end and smaller residential projects that maintain or enhance the character of their area, leading to more planning applications for such projects.

In column (3), we use the refusal rate for minor residential planning applications as the depen-

Table 5 – Homeownership and regulation: other dependent variables

Dependent variable:	Residential refusal rate, 5-year window (sd)	Log of residential planning applications	Minor residential refusal rate (sd)	Commerical refusal rate (sd)	Advertisement refusal rate (sd)
	(1)	(2)	(3)	(4)	(5)
Share owner-occupied housing (sd)	0.456***	0.079	0.596***	0.571***	-0.203
	(0.107)	(0.179)	(0.109)	(0.098)	(0.151)
Share council housing 1971×year dummies	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,598	1,681
Kleibergen-Paap F-statistic	555	568	568	325.5	563.9

Notes: In column (1), the residential planning refusal rate from $t-2$ to $t+2$ is used as the dependent variable. (sd) denotes variables standardized to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors, shown in parentheses, are adjusted for cross-sectional dependence based on a 10-year time window following Driscoll and Kraay (1998). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

dent variable and find a positive effect of homeownership on this alternative measure. Specifically, a one standard deviation increase in homeownership is associated with a 0.6 standard deviation increase in the refusal rate for minor applications. We also observe similar effects for the refusal rate of commercial planning applications in column (4). This is consistent with the notion that homeowners may also oppose commercial developments, which can trigger similar NIMBY-related concerns.

Finally, as a form of placebo analysis, we examine the refusal rate for advertisement consent—applications for permission to display certain types of advertisements, such as billboards or illuminated signs. Since these cases are unlikely to trigger strong NIMBY concerns, we do not expect homeownership rates to influence the refusal rate. Column (5) confirms this expectation, showing no significant effect.

Oster's bias-adjusted estimator. We also consider an alternative identification strategy by adopting Oster's (2019) bias-adjusted estimator, which addresses omitted variable bias by estimating how unobserved factors might be correlated to the treatment variable, similarly to observed factors. More specifically, by comparing changes in R^2 with shifts in the coefficient of interest, observable control variables can provide insights into unobservables, especially when the observables capture the most important factors correlated with the variable of interest.

This estimator requires setting two key parameters: $R^2_{\max}$, the hypothetical R^2 if all relevant variables were observed, and δ, representing the relative influence of unobserved versus observed

Table 6 – Homeownership and regulation: bias-adjusted estimates

Dependent variable:	Residential planning refusal rate (standardized)			
	Baseline bias-adjusted	5-year window	Proportional selection, δ=0.75	Unrelated control
	(1)	(2)	(3)	(4)
Share owner-occupied housing (sd)	1.535***	1.803***	0.867***	1.332***
	(0.494)	(0.578)	(0.171)	(0.404)
Controls	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓
$R^2_{\max}$	0.556	0.565	0.556	0.556
δ	1	1	0.75	1
Number of observations	1,682	1,682	1,682	1,682
R^2	0.461	0.461	0.461	0.461

Notes: The $R^2_{\max}$ is calculated by regressing the refusal rate in t on the average refusal rate in $t-1$ and $t+1$ in column (3). In column (2), $R^2_{\max}$ is obtained using a 5-year window by regressing the refusal rate in t on the average refusal rate in $t-2$, $t-1$, $t+1$, and $t+2$. (sd) denotes variables standardized to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. Standard errors are bootstrapped (500 replications) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

variables. With these two parameters, the adjusted estimate β_1^* accounts for potential omitted variable bias. Typically, $R^2_{\max}$ is set near the observed R^2. We determine $R^2_{\max}$ by using the R^2 from a regression of the refusal rate in t on the average refusal rate in $t-1$ and $t+1$, thereby purging 'true' variation in refusal rates from random variation that cannot be explained.[6] Oster (2019) suggests that $\delta = 1$ serves as a reasonable upper bound for proportional selection.

Table 6 presents the results. Using Oster's GMM estimator, which relies on changes in both the coefficient and R^2, the baseline bias-adjusted β_1^* in column (1) indicates a strong positive effect of homeownership on the refusal rate. The coefficient suggests that having a standard deviation increase in the homeownership rate raises the refusal rate by 1.5 standard deviations, which is somewhat stronger than the baseline IV estimate. A Hausman t-test does not indicate, however, that this estimate is statistically significantly stronger than the baseline IV estimate.[7]

In the remaining columns of Table 6, we assess the sensitivity of β_1^* under alternative assumptions. Column (2) uses a 5-year window to predict the refusal rate in t based on the average

[6] Based on experimental studies, Oster (2019) argues that a rule-of-thumb value for $R^2_{\max}$ can be obtained by multiplying the R^2 of a regression with controls by 1.3. In our case, this results in similar values for $R^2_{\max}$.

[7] Hausman's t-test is given by $(\hat{\beta}_1^* - \hat{\beta}_1^{\text{iv}})/\sqrt{\sigma^2_{\hat{\beta}_1^*} - \sigma^2_{\hat{\beta}_1^{\text{iv}}}}$, where $\hat{\beta}$ refers to the estimated coefficient and $\sigma_{\hat{\beta}}$ refers to its corresponding estimated standard error, see Wooldridge (2002).

refusal rate in the two years preceding and following t to estimate $R^2_{\max}$. This results in a slightly higher $R^2_{\max}$, leading to a somewhat stronger bias-adjusted estimate. Column (3) adopts a lower value for δ, suggesting that unobservables are relatively less important compared to observables; the resulting estimate is 0.87, which is close to the baseline IV estimate. In the final column, we treat the predicted employment shock as an 'unrelated' control and restrict the set of observable controls to weather and demographics. The fact that we obtain a similar coefficient indicates that our result does not critically depend on the choice of observable control variables.

There may be concerns about the sensitivity of Oster's estimator to chosen values for $R^2_{\max}$ and δ. A more philosophical critique is that it remains fundamentally unknown to what extent observables truly inform us about the nature of unobservable factors. Nevertheless, it is reassuring that a completely different set of assumptions, compared to those underlying the preferred IV estimates, leads to a similar conclusion: homeownership has a strong impact on regulatory restrictiveness.

3.4 Housing and developed land

We now examine the relationship between homeownership, new housing construction, and the expansion of developed land. Because regulatory constraints tend to be more stringent in areas with higher homeownership rates, one might expect new housing supply to be more limited. At the same time, national targets for housing construction (see e.g., Department for Communities and Local Government, 2007, 2017) suggest that local planning committees may have limited influence over the total number of new properties being constructed, but retain considerable discretion over where new development occurs. Given the prevalence of NIMBY concerns, we therefore anticipate that the impact of regulatory restrictions will be more pronounced on the share of land developed land than on the number of housing units constructed.

In Table 7, we first analyse dwelling construction, using as our dependent variable the change in the number of dwellings between t and $t+2$. Column (1) reports a small negative effect: a one standard deviation increase in the homeownership rate reduces the annual growth in housing units by approximately 0.6%. This effect may seem small, but it accumulates over time. After 10 years, areas with higher homeownership rates by one standard deviation exhibit roughly 6% ($(1-0.994^{10})\times 100\%$) fewer dwellings, a difference that is economically meaningful. The result remains robust when allowing for non-linear effects of the 1971 council housing share in column (2) and when including potentially endogenous controls in column (3).

Columns (4) to (6) replicate the analysis, but measure dwelling growth over a 10-year horizon, which necessitates the exclusion of the final wave of data. The estimated coefficients remain negative, although they are statistically insignificant. We interpret these results as providing a

Table 7 – Homeownership and new housing

Dependent variable:	Δ Log of number of dwellings, $t+2$			Δ Log of number of dwellings, $t+10$		
	Baseline IV	+Nonlinear shares	+Endogenous controls	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)	(4)	(5)	(6)
Share owner-occupied housing (sd)	-0.006**	-0.006**	-0.006**	-0.003	-0.003	-0.003
	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)	(0.003)
Residential planning applications (log)			0.002***			0.001***
			(0.000)			(0.000)
Male weekly earnings (log)			-0.000			-0.001
			(0.003)			(0.001)
Share council housing 1971×year dummies	✓	✓	✓	✓	✓	✓
(Share council housing 1971)^2×year dummies		✓	✓		✓	✓
Controls	✓	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,349	1,349	1,349
Kleibergen-Paap F-statistic	568	581.6	499.5	854.5	676.5	837.7

Notes: The dependent variable in columns (1)-(3) is the change in number of dwellings between $t+2$ and t and in columns (4)-(6) it is the change in number of dwellings between $t+10$ and t. This implies that the final wave of data from 2021 cannot be used in columns (4)–(6). (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

conservative lower bound for the effect of homeownership on dwelling construction.

In Table 8 we take the share of developed land as dependent variable. Having a consistent measure of developed land over time is not straightforward, which is why we have two different measures based on two very different data sources. The first is based on the Land Cover, and available since 1990. In columns (1)-(3) we show a strong negative effect of the homeownership rate on the share of developed land. The preferred specification suggests that a standard deviation increase in homeownership leads to an increase in the share of developed land of 0.27 standard deviations. The effect is robust to controlling for non-linear effects of council housing shares in 1971 and controlling for the number of planning applications and earnings. In other words, the likelihood of land being developed is substantially lower in areas with higher rates of homeownership.

Columns (4)-(6) take the share of urban land as dependent variable based on Li et al. (2021), available between 1980 and 2013. The results are virtually the same.

Table 8 – Homeownership and developed land

Dependent variable:	Share of developed land (standardised)			Share of urban land (standardised)		
	Baseline IV	+Nonlinear shares	+Endogenous controls	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)	(4)	(5)	(6)
Share owner-occupied housing (sd)	-0.271***	-0.269***	-0.267***	-0.257***	-0.266***	-0.279***
	(0.032)	(0.031)	(0.029)	(0.042)	(0.044)	(0.045)
Residential planning applications (log)			0.006***			0.015**
			(0.001)			(0.007)
Male weekly earnings (log)			0.010			0.125***
			(0.024)			(0.011)
Share council housing 1971×year dummies	✓	✓	✓	✓	✓	✓
(Share council housing 1971)^2×year dummies		✓	✓		✓	✓
Controls	✓	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Number of observations	1,370	1,370	1,370	1,349	1,349	1,349
Kleibergen-Paap F-statistic	140.5	160.7	157.6	854.5	676.5	837.7

Notes: The dependent variable in columns (1)–(3) is the share of developed land based on the Land Cover maps, which are available from 1990 onward. In columns (4)–(6), it is the share of urban land derived from Li et al. (2021), available up to 2013. (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

3.5 House prices and rents

Since a higher share of homeowners suggests greater regulatory restrictiveness, we expect this to raise house prices and rents by limiting the supply of new housing and developable land, and through associated amenity effects. Table 9 reports the results.

Column (1) shows large and statistically significant effects on house prices: a one standard deviation increase in the homeownership rate raises prices by approximately 40% ($(\exp(0.340) - 1) \times 100\%$). These estimates remain robust to the inclusion of non-linear controls for the 1971 council housing share, housing demand, and the number of planning applications.

Columns (4)–(6) shift the focus to rents and reveal substantially smaller effects. The point estimates are less than half those for house prices and are not statistically significant. We view the smaller rent effects as consistent with the idea that house prices are forward-looking, incorporating expectations about future housing supply and amenities, while rents respond mainly to conditions in the near term. Another interpretation is that homeowners may be effective at steering development away from their immediate surroundings, implying larger price increases around owner-occupied properties than around rented ones. A third interpretation

Table 9 – Homeownership, house prices, and rents

Dependent variable:	Log of house prices			Log of rents		
	Baseline IV	+Nonlinear shares	+Endogenous controls	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)	(4)	(5)	(6)
Share owner-occupied housing (sd)	0.340***	0.353***	0.346***	0.139	0.137	0.135
	(0.106)	(0.103)	(0.102)	(0.087)	(0.088)	(0.088)
Residential planning applications (log)			0.008***			0.026*
			(0.002)			(0.014)
Male weekly earnings (log)			0.129***			0.001
			(0.025)			(0.093)
Share council housing 1971×year dummies	✓	✓	✓	✓	✓	✓
(Share council housing 1971)^2×year dummies		✓	✓		✓	✓
Controls	✓	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,682	1,682	1,682
Kleibergen-Paap F-statistic	568	581.6	499.5	568	581.6	499.5

Notes: The dependent variable in columns (1)-(3) is the log of mean house prices and in columns (4)-(6) it is the log of mean rent. (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

is that homeowners have different preferences for open space and new developments (Asquith et al., 2023).

3.6 Mechanisms

We hypothesised earlier that homeowners are likely to exert influence on planning decisions through voting and direct influence. To investigate this further, we first turn our attention to voting outcomes in Table 10. We maintain the same set of specifications as in the baseline but replace the dependent variable with the share of Conservative seats in a local council. Remarkably, we observe robust and positive impacts of homeownership on the share of Conservative seats, irrespective of whether homeownership is instrumented for. The preferred specification in column (1) shows that a standard deviation increase in the homeownership rate increases the share of Conservative seats by 0.78 standard deviations. The overall positive effect carries significant implications in the context of the RtB policy: as RtB spurred homeownership, it likely bolstered support for the Conservative party.

Second, homeowners may seek to directly influence local planning decisions. Efforts may involve filing amendments and petitions, as well as mobilising neighbours and local communities

Table 10 – Homeownership, voting, and planning delay

Dependent variable:	Share Conservative seats (standardised)			Residential planning delay rate (standardised)		
	Baseline IV	+Nonlinear shares	+Endogenous controls	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)	(4)	(5)	(6)
Share owner-occupied housing (sd)	0.778**	0.803**	0.782**	0.836***	0.832**	0.829**
	(0.340)	(0.349)	(0.354)	(0.310)	(0.331)	(0.358)
Residential planning applications (log)			-0.067***			-0.134
			(0.020)			(0.082)
Male weekly earnings (log)			0.574***			0.325*
			(0.106)			(0.184)
Share council housing 1971×year dummies	✓	✓	✓	✓	✓	✓
(Share council housing 1971)^2×year dummies		✓	✓		✓	✓
Controls	✓	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,682	1,682	1,682
Kleibergen-Paap F-statistic	568	581.6	499.5	568	581.6	499.5

Notes: The dependent variable in columns (1)-(3) is the share of Conservative seats in local authorities and in columns (4)-(6) it is the the residential planning delay rate, which is defined as the share of planning decisions that took more than 8 weeks. (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. We instrument the share of owner-occupiers with the log of the predicted share of owner-occupied housing based on the total number of right-to-buy sales in each year and the initial share of council housing in 1971. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

to protest against specific new developments. Regardless of the outcome of such activities, they inevitably prolong the planning process. To assess this impact, we utilise ancillary data on the share of planning decisions that took longer than 8 weeks. The instrumental variable specifications demonstrate positive and statistically significant effects. The preferred specification in column (4) suggests that a standard deviation increase in the homeownership rate increases the delay rate by 0.84 standard deviations, indicating a non-negligible impact.

4 A dynamic spatial equilibrium model

In this section, we present a dynamic quantitative model of England. It builds on the model developed in Greaney et al. (2025). The key ingredients of the model are (1) many locations, each of which can serve as residence and workplace, (2) commuting between residence and workplace, (3) forward-looking agents, (4) tenure choice with illiquid owner-occupied housing, (5) endogenous floorspace supply, rents, and prices, as well as (6) endogenous land use regulation. The model is set in continuous time.

5 Conclusions

This paper provides robust reduced-form evidence that higher local homeownership rates lead to more restrictive land use regulation in England. Exploiting exogenous variation from the Right-to-Buy policy, we show that increased homeownership raises planning refusal rates and decision delays, reduces housing construction and land development, and capitalises into higher house prices, with much weaker and insignificant effects on rents. These findings are consistent with a homevoter mechanism operating through both electoral and procedural channels.

To interpret these patterns and assess their welfare implications, we develop a dynamic spatial equilibrium model with homeowners and renters, endogenous development, and a durable housing stock. The model rationalises the reduced-form results by showing how higher homeownership curtails new development and raises prices, while potentially generating amenity gains that are valued asymmetrically across tenures. Counterfactual simulations assess the welfare costs of NIMBYism…

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Appendix A Theory

[...]

Appendix B Data

B.1 Annual panel dataset

In addition to the decennial panel dataset, we construct an annual panel spanning 1979–2021. Most variables are available on an annual basis, with the exception of housing unit counts by tenure. The share of owner-occupied housing is estimated using a combination of census data, construction statistics by tenure from 1980 to 2000, and data from the DLHC, which provides interpolated dwelling and tenure statistics from 2001 onward. Since Census data are the most reliable, we ensure that trends in non-Census years align with those in Census years so that the dwelling counts by tenure remain consistent with observed Census figures.

Table B1 presents descriptive statistics for the annual panel dataset. The summary statistics do not show substantial differences from those in the decennial panel.

B.2 Additional descriptive statistics

Figure B1 illustrates the trends in planning refusal rates and delay rates for major residential projects (exceeding ten units) in England. The refusal rate follows a cyclical pattern, increasing during periods of higher housing demand. The delay rate remains consistently high, suggesting persistent inefficiencies or intentional slowdowns in the planning process and was only reduced

Table B1 – Descriptive statistics for the annual panel dataset

	Mean	Standard deviation	Min	Max
	(1)	(2)	(3)	(4)
Residential planning refusal rate (%)	23.82	17.07	0	100
Residential planning delay rate (%)	63.35	20.11	0	100
Residential planning applications	22.00	17.01	1	221
Share owner-occupied housing (%)	68.18	10.99	6.118	92.21
Share council housing (%)	12.99	10.45	0	80.81
Share housing association units (%)	5.188	4.039	0	26.46
Share private rent (%)	13.64	6.010	1.309	47.83
Number of dwellings	55,833	37,716	662	423,529
Mean house price (£)	148,419	142,314	10,937	2.081e+06
Mean annual rent (£)	7,780	5,622	505.9	88,822
Share of Conservative seats (%)	41.12	24.84	0	100
Share of Labour seats (%)	31.54	28.40	0	100
Share of liberal-democrat seats (%)	15.48	17.01	0	94.23
Share of other seats (%)	11.85	16.76	0	100
Male weekly earnings (£)	446.0	212.2	0	1,793
Predicted employment	73,319	57,511	1,182	621,264
Predicted share of owner-occupied housing (%)	76.53	13.19	4.815	95.05

Notes: The number of local authorities included is 354. The residential planning delay rate is defined as the share of planning decisions that took more than 8 weeks. We adopt the pre-2009 definition of Local Authorities (LAs). Subsequent to 2009, the number of LAs has been reduced to 326. Where necessary, we adjust post-2009 values to align with the LAs in 2009, considering weights based on the 2009 population. Here, we do not report the weather variables, including minimum temperature, maximum temperature, sunshine hours, and rainfall. Those values are available upon request.

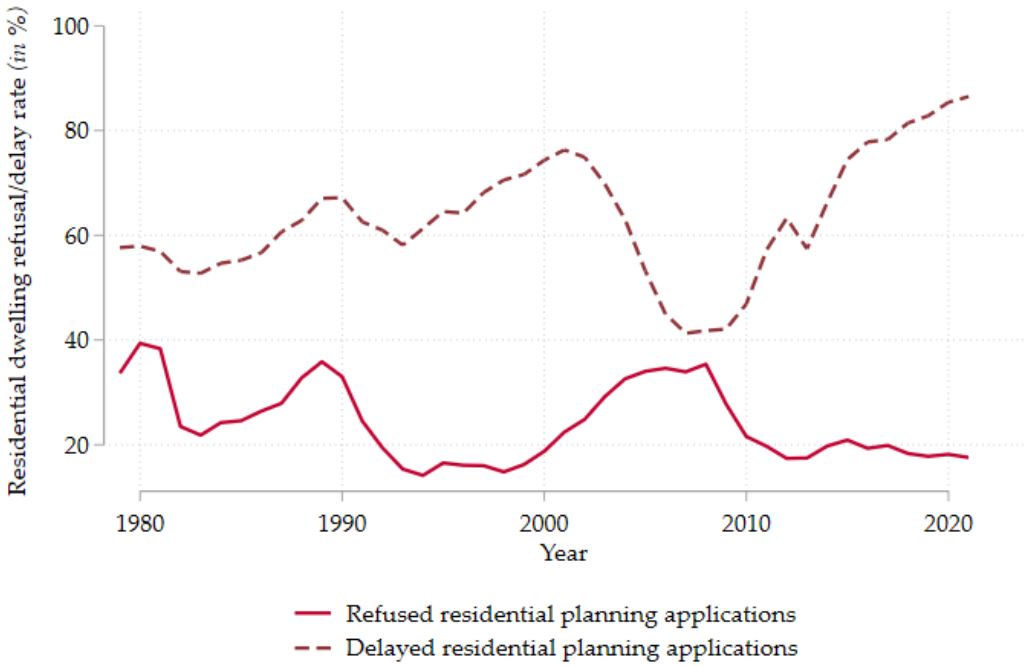


Figure B1 – Regulatory restrictiveness over time

Notes: We plot here the refusal rate and delay rate for major residential planning applications in England, which exceed 10 properties. The delay rate is calculated as the percentage of planning applications that require more than 8 weeks to be processed and reach a decision.

post 2000 in response to endeavours to reduce planning delay.

Figure B2 presents the share of local council seats held by the Conservative and Labour parties over time. A notable electoral cycle emerged in the mid-1990s, when the Labour Party dominated local councils. In contrast, over the past decade, the distribution of seats has been relatively balanced, with both parties each holding approximately 40% of the seats.

Two histograms are reported in Figure B3, one for regulatory restrictiveness and the other for owner-occupied housing. The distribution of refusal rates across local authorities reveals that there is considerable variation in planning application rejection rates. Similarly, the home-ownership histogram shows that homeownership rates hover around 65%, although there is considerable variation across space and time.

Figure B4 tracks the predicted share of homeownership over time, as defined in (3), based on the instrument derived from the Right-to-Buy (RtB) policy. The upward trend illustrates the impact of RtB on local homeownership rates, supporting the first-stage results that show a strong and robust relationship between the predicted share of homeownership and actual homeownership levels.

Figure B5 consists of two panels. In Figure B5a we present spatial the distribution of council

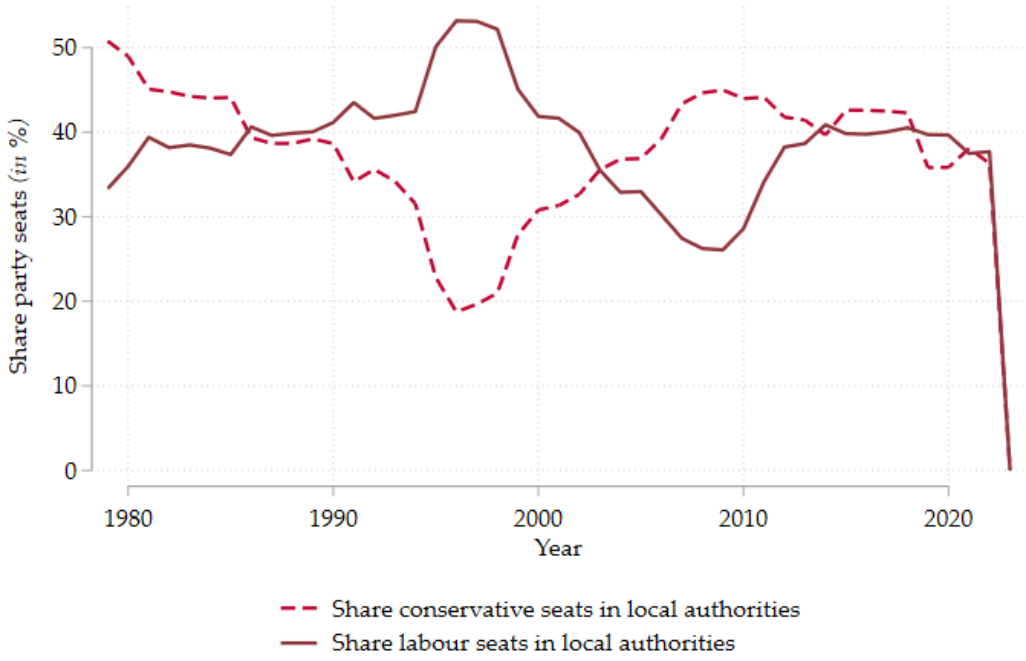
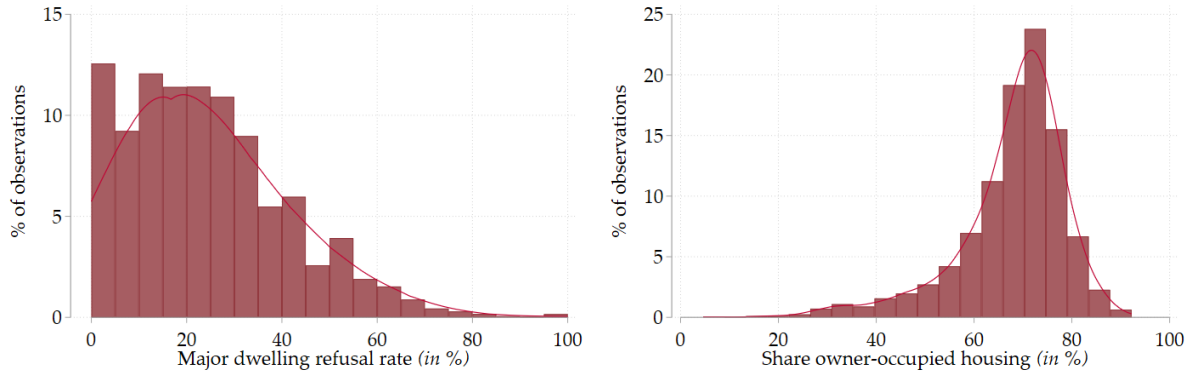


Figure B2 – Local election outcomes by party

Notes: We plot here the share of overall seats in local governments for the two main parties: the Conservative party and the Labour party.



(a) Regulatory restrictiveness (b) Owner-occupied housing

Figure B3 – Histograms for key variables

Notes: These are histograms based on local-authority level data for 5 waves of data: 1981, 1991, 2001, 2011, and 2021.

housing shares across local authorities in 1971. Areas with a high initial share of council housing were more affected by the Right-to-Buy (RtB) policy, leading to significant shifts in homeownership rates over time. Figure B5b illustrates the spatial distribution of average house prices across local authorities from 1974 to 2021. The map reveals that house prices tend to be higher in Greater London and the Southeast, whereas they are comparatively lower in the North of England.

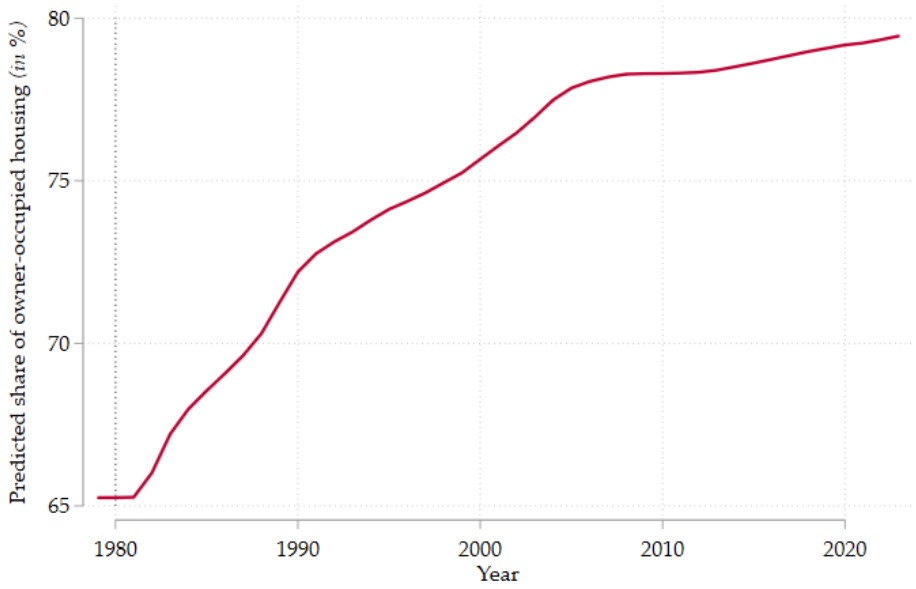
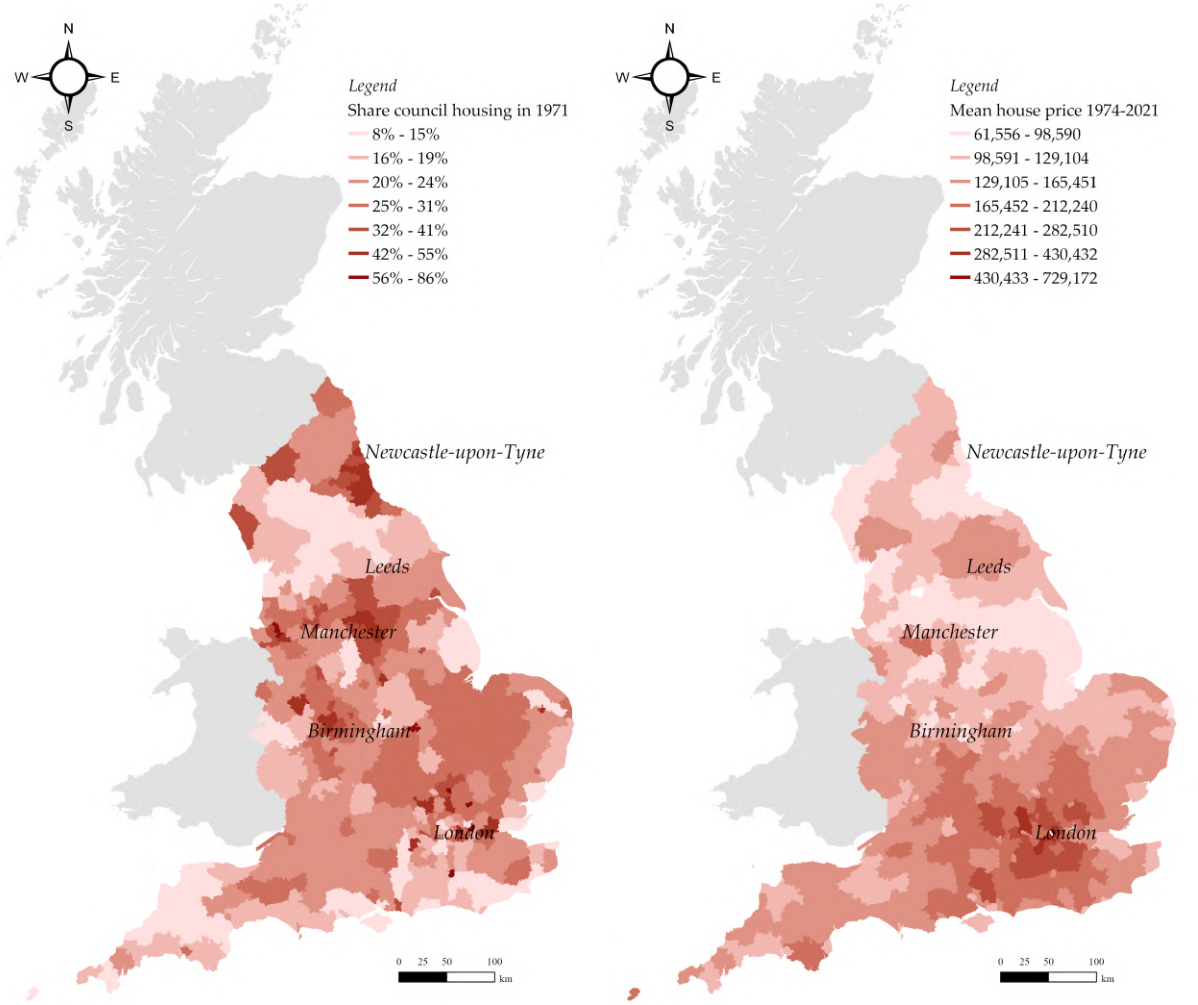


Figure B4 – Predicted share of homeownership over time

Notes: The measure depicted here is calculated as $\left(n^o_{i1971} \times (n^o_{1971} + \text{rtb}_t)/n^o_{1971}\right)/\left(n^o_{i1971} \times (n^o_{1971} + \text{rtb}_t)/n^o_{1971} + n^c_{i1971} \times (n^c_{1971} - \text{rtb}_t)/n^c_{1971}\right)$, where n^o_{i1971} and n^c_{i1971} are, respectively, the number of owner-occupied and council housing units in 1971 in i, n^o_{1971} are the total number of owner-occupied dwellings in 1971 in England, and rtb_t are the cumulative number of RtB sales since the policy's inception in 1980.



(a) Share council housing in 1971 (b) Mean house prices, 1974-2021

Figure B5 – Council housing in 1971 and house prices across space

Notes: These are histograms based on local-authority level data.

Appendix C Other reduced-form results

C.1 First-stage results

We present the first-stage results in Table C1. We regress the homeownership rate on the log of the predicted share of owner-occupied housing, as defined in (3). The results demonstrate strong and robust positive effects across various specifications. We find that doubling the predicted share of owner-occupied housing increases the share of owner-occupied housing by about 4.5 standard deviations ($(\ln 2 - \ln 1) \times 6.414$). The effect is robust across specifications.

Table C1 – Homeownership and regulation: first-stage results

Dependent variable:	Share owner-occupied housing (standardised)		
	Baseline IV	+Nonlinear shares	+Endogenous controls
	(1)	(2)	(3)
Predicted share of owner-occupied housing (log)	6.414***	6.338***	6.271***
	(0.239)	(0.234)	(0.249)
Residential planning applications (log)			0.023***
			(0.007)
Male weekly earnings (log)			0.179
			(0.112)
Share council housing 1971×year dummies	✓	✓	✓
(Share council housing 1971)^2×year dummies		✓	✓
Controls	✓	✓	✓
Local authority fixed effects	✓	✓	✓
Year fixed effects	✓	✓	✓
Number of observations	1,682	1,682	1,682
R^2	0.976	0.971	0.971

Notes: (sd) denotes variables standardised to mean zero and unit standard deviation. Controls include the predicted employment, which is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. They further include the minimum temperature, maximum temperature, sunshine hours, rainfall, the share of people younger than 30, the share of people 30-64 years, the share of people older than 64, the share of people with (permanent) disabilities, and the share of people with a college education. Standard errors are adjusted for cross-sectional dependence based on a time window of 10 years following Driscoll and Kraay (1998) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.