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

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Abstract — We study the effects of endogenous land use regulation on welfare and sorting across space. Utilizing panel data from England and employing the Thatcherian right-to-buy policy as an exogenous shifter in local homeownership, we first establish that changes in homeownership lead to an increased share of refused planning applications, supporting the 'homevoter hypothesis'. Higher levels of homeownership are also associated with greater house price appreciation. Next, we set up a dynamic spatial equilibrium model in which increased regulatory strictness results in decreased construction activity but possibly elevated amenity levels. Our objective is to conduct counterfactual analyses by eliminating 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 decaded years, 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 shows 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 significant role in cities with fewer geographical constraints, such as London and Paris. As a result, the importance of endogenous supply constraints, such as land use regulations, becomes critical.

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

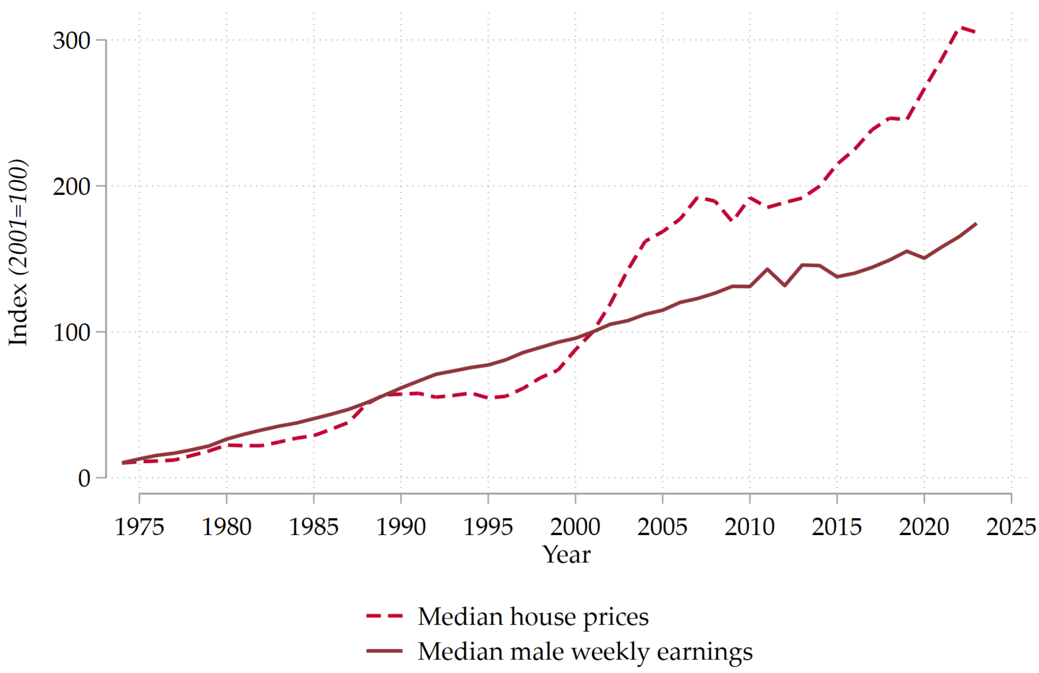


FIGURE 1 – HOUSE PRICES AND EARNINGS OVER TIME

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., 2005*b*; 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 especially true when homeowners form the majority in a given area. By supporting more restrictive policies, they aim to preserve or increase their property values by preventing declines in local amenities (such as the loss of open space to new development) and by limiting the supply of housing. In doing so, homeowners often prioritize 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. 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), we use the number of refused planning applications of major residential projects as our main indicator of regulatory restrictiveness. This provides us with 40 years (1981-2021) of data.

Because land use regulation is highly endogenous with respect to whether an LA has a majority of homeowners, we exploit 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 all specifications, we incorporate LA fixed effects to examine the impact of changes in homeownership on changes in the refusal rate. Thus, our identifying variation hinges on the premise that areas with a high concentration of council housing are prone to experiencing significant unexpected shifts in homeownership rates, potentially influencing the stringency of land use regulations. We provide a battery of robustness checks aiming to show that controlling for changes in local housing demand does not fundamentally alter the estimates. Applying this strategy, we find that having a majority of homeowners increases the share of refused planning applications by about 1 standard deviation, which is a non-negligible effect. We also show a positive effect of increased homeownership on house price appreciation: having a majority of homeowners increases house price change by 5% per annum. These effects hold up under extensive sensitivity analyses and alternative identification strategies. We highlight two key robustness checks here. First, given that initial homeownership shares in 1971 may be viewed as endogenous (see Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022) we limit ourselves to LAs that had homeownership shares just below 50%, leading to similar results. Second, instead of relying on an instrumental variables approach, we use Oster's (2019) procedure to adjust OLS estimates for potential omitted variable bias, yielding very similar results.

We hypothesize that the effect of homeownership on regulation comes about via two channels. First, residents can exercise their influence by voting. Elected councilors have the authority to shape and approve planning policies and regulations 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). Hence, we expect that a majority of homeownership increases local votes for the Conservative party. Second, concerned homeowners can try to directly influence planning decisions. This could involve meeting with councilors as well as submitting petitions and amendments. This will inevitably delay the planning process.

We find evidence in support of these two channels. Specifically, having a majority of homeowners increases the proportion of Conservative seats increases by approximately 1.3 standard deviations. To gauge the impact of direct influence, we analyze the proportion of planning projects experiencing delays, defined as taking more than 13 weeks for a decision. Albeit suggestive, the results suggest that a having a majority of homeowners increases the delay rate of planning decisions by about 1 standard deviation.

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 increases homeownership levels as well as higher levels of development are associated with increased restrictiveness in the US. 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 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 centered 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., 2005*a*,*b*; Green et al., 2005; Ihlanfeldt, 2007; Hilber and Vermeulen, 2016; Severen and Plantinga, 2018). 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 emphasize 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 labor 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)

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 above papers typically take an omnibus measure of land use constraints and 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 endeavors 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 neighborhoods (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 that increased homeownership shares may enhance local amenity levels.

The remainder of this paper is structured as follows. Section 2 outlines the 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 descriptives

2.1 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 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 utilize data on the proportion of 'delayed' major dwelling planning applications, defined as the percentage of applications that exceed a duration of 13 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

[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. 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.

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.

Fifth, we obtain data on house prices, which are derived from transaction price data from the COUNCIL OF MORTGAGE LENDERS (1974-1995) and the LAND REGISTRY (1995-2021). We create real prices by deflating house prices by the Retail Price Index. We further 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 data and methodologies employed, we refer to Hilber and Vermeulen (2016). We also construct a labor 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.

2.2 *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,

[2]For the LA Redcar and Cleveland we do not observe the share of Labor 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 have 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.

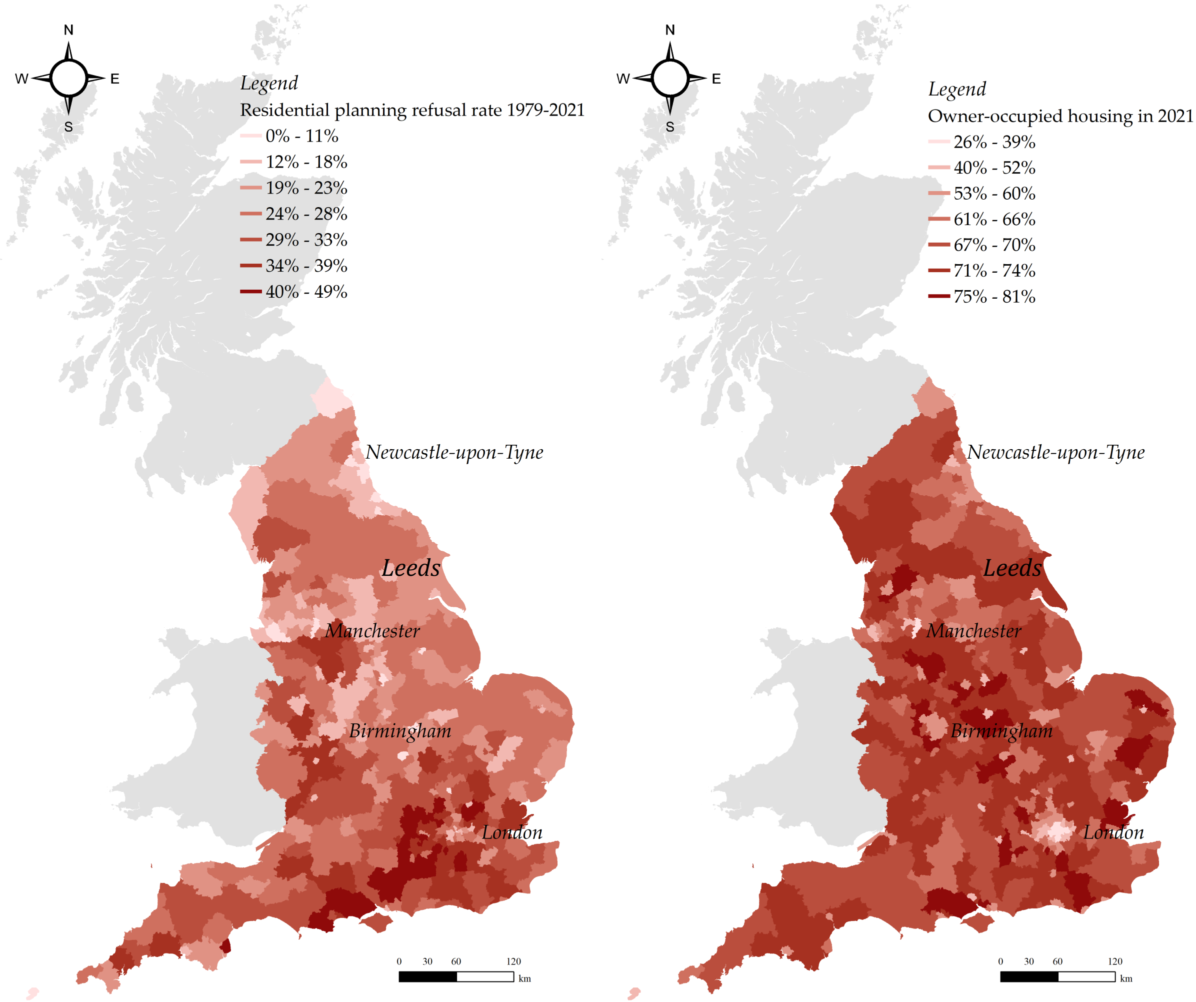
TABLE 1 – DESCRIPTIVE STATISTICS

	Mean	Standard deviation	Min	Max
	(1)	(2)	(3)	(4)
Residential planning refusal rate *(%)*	24.33	17.42	0	100
Residential planning delay rate *(%)*	80.66	19.91	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
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
Mean house price *(£)*	160,032	158,200	15,183	1,982,000
Male weekly earnings *(£)*	468.1	238.4	89.54	1,793
Predicted employment	62,694	51,134	1,663	551,706
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.

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 follow 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 about 75%, and exhibits less cyclicality compared to the share of refused planning 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%. It is important to 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%.



(A) RESIDENTIAL PLANNING REFUSAL RATE (B) SHARE OWNER-OCCUPIED HOUSING

FIGURE 2 – SPATIAL VARIATION

Notes: These are histograms 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.

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 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

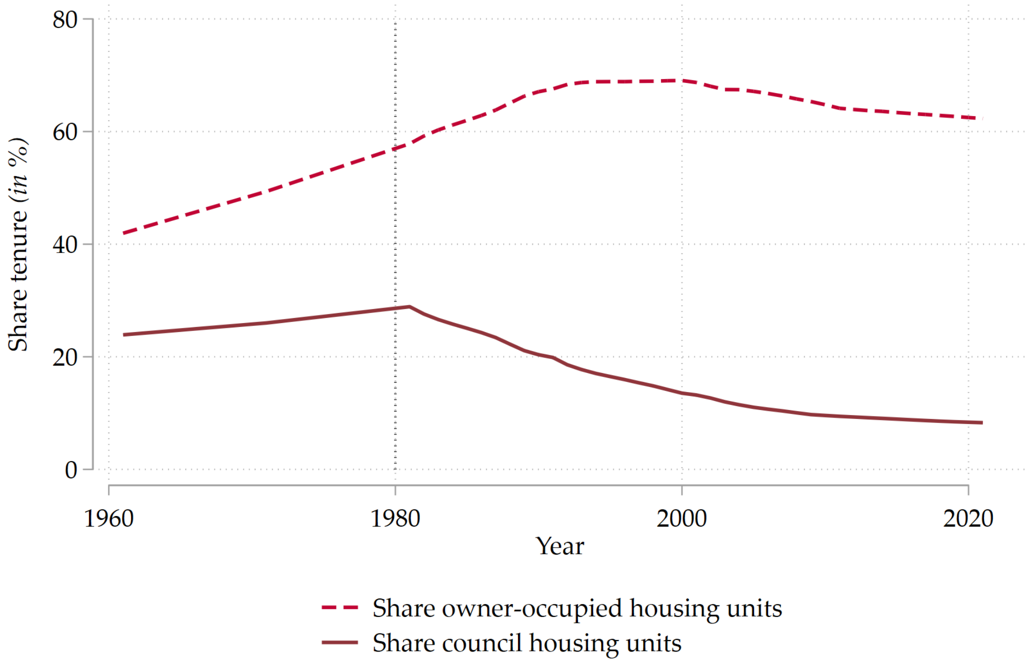


FIGURE 3 – HOUSING TENURE OVER TIME

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 DCLG, 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.

of electoral cycles, along with histograms depicting the key variables.

3 Reduced-form evidence

3.1 Methodology

Baseline estimation. Our main focus here is to demonstrate the favorable impact of increased levels of homeownership on regulatory restrictiveness, as proxied by the residential planning refusal rate. Additionally, we seek to illustrate that this effect occurs through two channels: homeowners voting for parties that endeavor 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 m_{it} is a dummy variable indicating whether a LA had a majority of owner-occupied housing in t. We then estimate:

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

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

This equation is subject to endogeneity 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. New developments are more frequently rental projects.[4] Due to the cyclical nature of refusal rates (recall Figure B1), 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 m_{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 utilizing data from the Department of Local Housing and Communities (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

[4] In our data, the cross-sectional correlation between the dummy variable indicating whether an LA has a homeowner majority and the Bartik demand shock is -0.282. Conditional on LA and year fixed effects, the (partial) correlation is -0.051.

exhibited a high proportion of council housing. Let us define:

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

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:

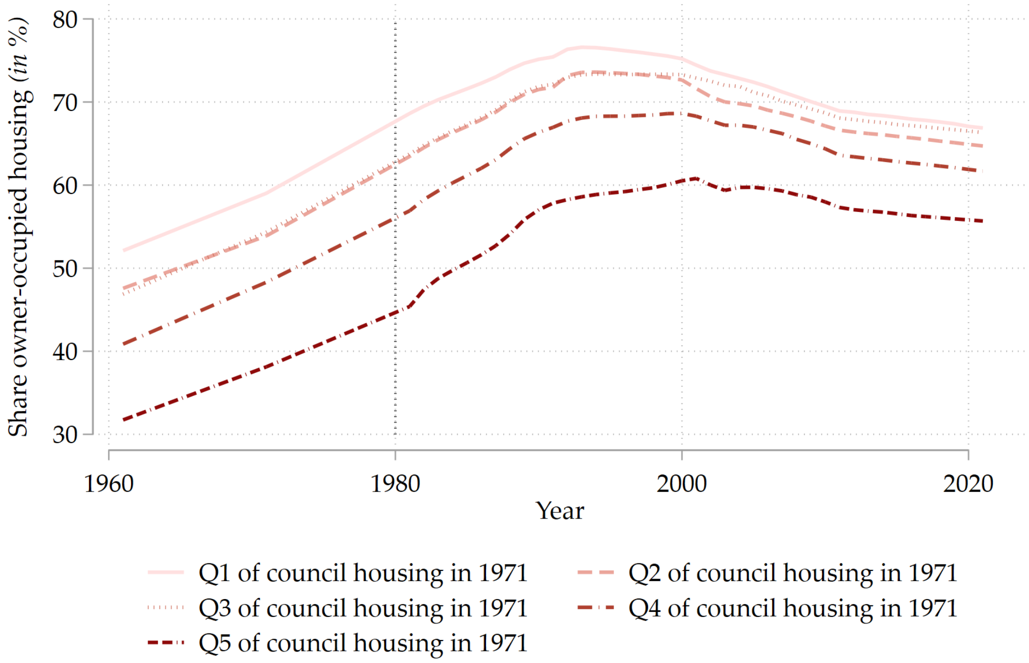
$$z_{it} = \frac{\hat{n}^o_{it}}{\hat{n}^o_{it} + \hat{n}^c_{it}}, \tag{3}$$

so z_{it} represents 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, z_{it} increases from about 65% in 1980 to about 80% in 2021.

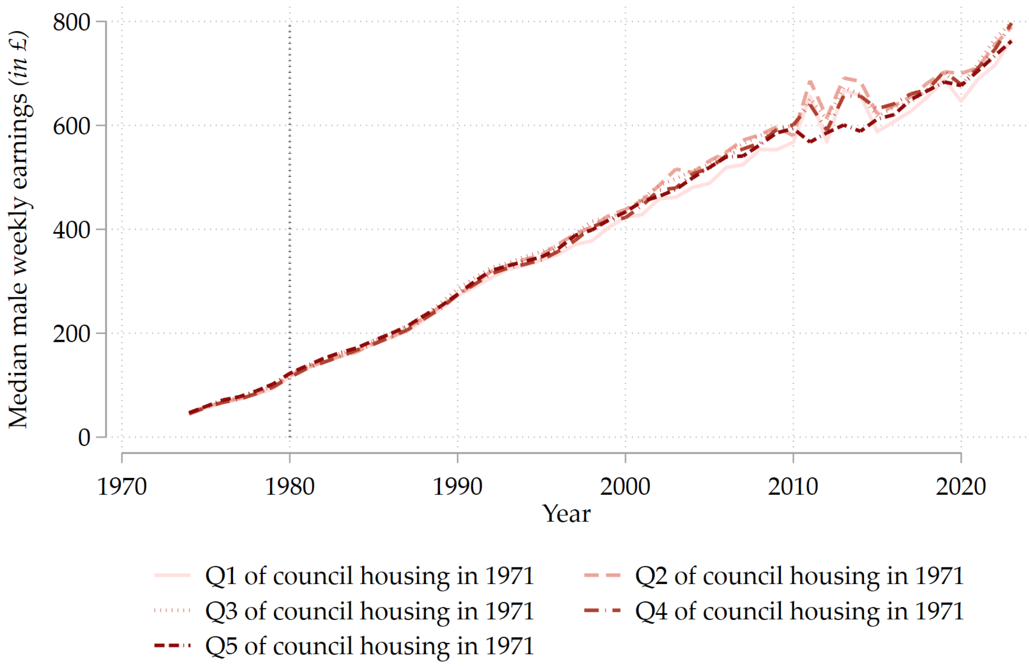
The primary concern with the instrument is that LAs with high initial shares of council housing may experience lower housing demand, leading local planning authorities to approve nearly all construction projects in the following years. This suggests that the initial shares are correlated with changes in refusal rates. This relates to a more general criticism of shift-share instruments that the identification comes from the initial shares in housing tenure, which may be endogenous (Goldsmith-Pinkham et al., 2020).[5] One thing one can do is inspect pre-trends in initial shares. We show in Figure 4 supporting the exogeneity of initial shares 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.

To address any residual endogeneity concerns, our preferred specifications include controls for local earnings or 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 from aggregate changes in the industry structure in England. Additionally, we focus on LAs with similar homeownership shares in 1971. Specifically, in some specifications, we include LAs with ownership shares

[5]There might be 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 -0.1. 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.

between 25% and 75% in 1971. In a more restrictive specification, we narrow this range to 45% to 55%.

We also estimate a specification where we control for LA-specific linear trends, which inevitably leads to higher standard errors. Further, we conduct regression analyses using the annual panel data, where we focus on the years around the enactment of the RtB policy. This focus is warranted as the majority of the variation in the instrument stems from earlier years, given that a significant portion of RtB sales occurred in the initial years following the policy's implementation. 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 house price changes. Subsequently, we examine the effects of homeownership on house price changes. Hilber and Vermeulen (2016) demonstrate that regulation prompts home price appreciation in more tightly regulated LAs. Consequently, we anticipate that, because a majority of homeowners will spur greater regulatory restrictiveness, this ultimately results in increased home price appreciation. We then proceed to estimate:

$$\frac{p_{it+5}/p_{it}}{5} = \beta_1 m_{it} + \beta_2 x_{it} + \mu_i + \mu_t + \epsilon_{it}, \tag{4}$$

where p_{it+5} is the mean house price in $t + 5$. Once more, we instrument for m_{it} with z_{it}.

We acknowledge that the estimated coefficient β_1 in the above specification might also capture favorable amenity effects (such as the preservation of green space) and potentially, positive externalities associated with increased homeownership. In our spatial equilibrium model, we aim to disentangle the supply effects and the amenity effects 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:

$$\{v_{it}, \ell_{it}\} = \beta_1 m_{it} + \beta_2 x_{it} + \mu_i + \mu_t + \epsilon_{it}, \tag{5}$$

Table 2 – Owner-occupied housing and regulation

Dependent variable:	*Residential planning refusal rate (standardized)*					
	Naïve OLS	Baseline IV	+Control variables	+Earnings control	+Demand shock	Homeowners in 1971, 25-75%
	(1)	(2)	(3)	(4)	(5)	(6)
Majority of homeowners	0.044	0.979**	0.865**	0.914***	0.973***	1.088***
	(0.100)	(0.390)	(0.426)	(0.199)	(0.261)	(0.387)
Residential planning applications *(log)*	0.048	0.044	0.038	0.000	0.002	0.025
	(0.056)	(0.054)	(0.049)	(0.057)	(0.058)	(0.057)
Male weekly earnings *(log)*				0.694***		
				(0.157)		
Predicted employment *(log)*					0.459	0.103
					(0.291)	(0.241)
Weather and demographic controls			✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Number of observations	1,682	1,682	1,682	1,682	1,682	1,594
R^2	0.450					
Kleibergen-Paap F-statistic		3.556	3.674	6.938	6.470	5.556

Notes: Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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 (2)-(6) we instrument for dummy indicating whether an LA has am majority of owner-occupiers with 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. The predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 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 13 weeks. We instrument for m_{it} using (3).

3.2 *Baseline results*

The baseline results are presented in Table 2. The OLS specification reveals a a small yet statistically insignificant coefficient for whether a LA has a majority of homeowners. We already indicated that naïve OLS estimates are likely underestimated.

Hence, in column (2) we employ an instrumental variable approach, using the predicted share of owner-occupied housing based on right-to-buy sales as an instrument. In Appendix C.1, we show that the predicted share of owner-occupied housing has a robust and positive effect on whether there is a majority of owner-occupied housing within an LA. However, the Kleibergen-Paap F-statistic indicates that our instrument maybe somewhat weak. We come back to this issue in the more comprehensive specifications.

Going back to Table 2, Column (2) reveals a robust and statistically significant positive effect of having a majority of owner-occupiers on regulation: having a majority of homeowners increases the refusal rate by 0.98 standard deviations. Interestingly, the number of planning applications exhibits no discernible impact on the refusal rate. In other words, there is no evidence to suggest that a high volume of applications overwhelms LAs, leading them to indiscriminately reject planning proposals.

In Column (3), we introduce a range of weather and demographic controls, including the share of people holding a degree. If regulation diminishes housing affordability, consequently fostering segregation based on educational attainment and income levels, controlling for changes in education levels could be important. However, the coefficient pertaining to the share of owner-occupied housing remains largely unchanged.

As demonstrated earlier, the refusal rate displays a pronounced cyclical pattern (recall Figure B1) and seems to be influenced by local housing demand. In Column (4), we employ the average weekly earnings of males as a proxy for local housing demand. While there is indeed a strong association between demand and the refusal rate (a doubling of earnings is associated with almost a 0.5 standard deviation increase in the refusal rate), the coefficient of interest remains essentially unaltered. Concerns regarding the endogeneity of earnings may arise; therefore, in our preferred specification in column (5), we control for a demand shock derived based on 1971 local industry shares. Notably, the predicted demand shock does not yield a statistically significant effect on the refusal rate, although this is mainly due to large standard errors. The impact of owner-occupied housing remains largely unaffected. Going back to the potential issue of weak instruments, based on critical values provided by Stock and Yogo (2005), any bias may be maximally 20-25% of the estimate, which means that our conclusions do not qualitatively change.

To further address concerns that the initial share of owner-occupied housing may correlate with unobserved regulatory trends, column (6) includes only LAs with homeownership shares between 25 and 75% in 1971. LAs with homeownership rates just below 50% in 1971 likely contribute the most to the relevant identifying variation, as these areas are the most likely to reach a homeowner majority. LAs that already had a small homeowner majority serve as suitable control areas. The coefficient indicates that the main effect remains largely unchanged.

3.3 Robustness

Main robustness analyses. Table 3 presents several robustness checks. In Column (1), we only select areas with homeownership rates between 45% and 55% to address any residual concerns that initial shares of owner-occupied housing are correlated to regulatory trends. We think it is reassuring that the point estimate hardly changes, although the standard errors are too large to draw very strong conclusions.

In Column (2), we adopt an alternative approach to control for trends in demand by incorporating 354 linear LA-specific trends. In this specification, having a majority of homeowners increases regulatory restrictiveness by 1.12 standard deviations, which is not materially different from the baseline specification.

Column (3) introduces two potentially endogenous controls potentially 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 instrumentation strategy.

In Column (4), 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 yet is somewhat smaller than in the preferred specification reported in Table 2.

In columns (5) and (6), we explore alternative measures for homeownership. Column (5) uses the share of homeowners within a Local Authority (LA). Unsurprisingly, the first-stage F-statistic is much higher here, as a continuous variable tends to yield a stronger first stage. The second-stage results indicate that a one standard deviation increase in homeownership raises the refusal rate by 0.19 standard deviations. In column (6), we replace our baseline dummy variable, which indicates a simple majority of homeowners, with a dummy for a qualified majority (*i.e.*, over 60%). This adjustment again strengthens the first stage, though the second-stage coefficient is slightly lower. We believe this may be due to the introduction of measurement error, as we would already expect regulation effects in LAs where homeowner shares are between 50% and 60%.

In the next 2 columns of Table 3, we switch to an annual panel dataset (details are provided in Appendix B.1). Column (7) examines a 20-year panel covering 1979–1999. This approach provides many more observations, a strong first stage, and a second-stage coefficient that closely aligns with the baseline coefficient. Column (8) narrows the focus to the years surrounding the policy implementation, where most of our identifying variation comes from; it includes one year before and 5 years after the Right to Buy (RtB) policy.[6] Here, we again find a positive effect of a majority owner-occupied housing on refusal rates, supporting the robustness of this result.

In the final columns of Table 3, we substitute the dependent variable with alternative measures

[6]The year 1980 was excluded, as it was the implementation year, likely introducing some uncertainty for prospective homeowners.

TABLE 3 – OWNER-OCCUPIED HOUSING AND REGULATION: ROBUSTNESS

Dependent variable:	*Residential planning refusal rate (standardized)*								*Log number of res. applications*	*Minor res. applications (sd)*	*Commercial applications (sd)*
	Decennial panel						Annual panel		Decennial panel		
	Homeowners in 1971, 45-55%	Linear LA trends	House prices	5-year window	Share homeowners	Qualified majority	1979-1999	1979-1985	Planning applications	Minor applications	Commerical applications
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Majority of homeowners	0.929	1.122**	0.958***	0.530***			0.960***	1.071**	0.248	-0.130	0.316
	(1.346)	(0.542)	(0.285)	(0.197)			(0.281)	(0.468)	(0.518)	(0.617)	(0.688)
Share owner-occupied housing *(sd)*					0.189**						
					(0.076)						
Qualified majority of homeowners						0.633**					
						(0.290)					
Residential planning applications *(log)*	-0.077**	0.047	-0.005	0.003	0.002	0.139***	0.088**	0.129**			
	(0.038)	(0.039)	(0.059)	(0.011)	(0.064)	(0.049)	(0.044)	(0.054)			
Minor residential planning applications *(log)*										0.006	
										(0.049)	
Commercial planning applications *(log)*											0.039
											(0.037)
Dwellings *(log)*			-0.168								
			(0.475)								
Mean house price *(log)*			0.333								
			(0.309)								
Predicted employment *(log)*	0.243	3.911**	0.093	-0.287	0.184	6.737***	0.917	8.156***	-0.980***	-0.224	-2.100***
	(0.898)	(1.638)	(0.582)	(0.769)	(0.252)	(1.387)	(1.034)	(2.134)	(0.238)	(0.194)	(0.382)
Weather controls	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Demographic controls	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Local authority × year trends		✓									
Number of observations	560	1,682	1,682	1,349	1,682	1,571	6,243	1,571	1,682	1,682	1,598
Kleibergen-Paap F-statistic	123.3	131.2	6.259	9.249	891.2	180.2	155.2	142.3	6.324	6.018	6.395

Notes: In column (4) we take the refusal rate from $t-2$ to $t+2$ as the dependent variable. *(sd)* refers to variables standardized with mean zero and unit standard deviation. A qualified majority is defined as being above 60%. Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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. We instrument for the share of owner-occupied housing with 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. The predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 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. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

of regulatory restrictiveness. Column (9) examines the effect of a homeownership majority on the log number of planning applications, finding no strong positive effect, which is reassuring. In column (10), we use minor residential planning applications as the dependent variable, and again find no effect of homeownership majority. This result is intuitive: homeowners likely do not influence approvals for extensions or attics, as they may also wish to make such improvements in the future. Finally, column (11) uses the number of commercial applications as the dependent variable, with no positive effect of homeownership detected here either. Homeowners may refrain from opposing commercial developments, as these projects create jobs and enhance amenities, potentially raising property values, despite potential NIMBY concerns.

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 by 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 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' varation in refusal rates from random variation that cannot be explained.[7] Oster (2019) suggests that $\delta = 1$ serves as a reasonable upper bound for proportional selection.

Table 4 presents the results. In column (1), we repeat the naïve OLS regression without additional controls, except for the number of residential planning applications, along with local authority and year fixed effects. In column (2), we include weather and demographic controls as well as the Bartik employment shock, which causes the coefficient to nearly quintuple. While the increase in R^2 appears modest, it is worth noting that at most 55% of the variation in the refusal rate can be explained. Using Oster's GMM estimator, which relies on changes in both the coefficient and R^2, the baseline bias-adjusted β_1^* in column (3) indicates a strong positive effect of a homeowner majority on the refusal rate. The coefficient suggests that a homeowner majority raises the refusal rate by 1.3 standard deviations, which is comparable to, but slightly

[7] 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 $R^2_{\max}$.

TABLE 4 – OWNER-OCCUPIED HOUSING AND REGULATION: BIAS-ADJUSTED ESTIMATES

Dependent variable:	*Residential planning refusal rate (standardized)*					
	Naïve OLS	+Control variables	Baseline bias-adjusted	5-year window	Proportional selection, $\delta = 0.75$	Add unrelated controls
	(1)	(2)	(3)	(4)	(5)	(6)
Majority of homeowners	0.044	0.215	1.296***	1.511***	0.927***	0.939***
	(0.163)	(0.161)	(0.416)	(0.500)	(0.254)	(0.332)
Residential planning applications	✓	✓	✓	✓	✓	✓
Predicted employment		✓	✓	✓	✓	✓
Weather and demographic 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	1,682	1,682
R^2	0.450	0.462	0.462	0.462	0.462	0.462

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 columns (3), (5), and (6). In column (4), $R^2_{\max}$ 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 $t+2$. Column (5) adjusts δ to 0.75, giving unobserved variables somewhat less weight. Column (6) includes the log of the number of planning applications and predicted employment as additional control variables, along with local authority and year fixed effects. Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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 predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 1971. Standard errors are bootstrapped (500 replications) and in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

stronger than, the baseline IV estimate.

In the remaining columns of Table 4, we assess the sensitivity of β_1^* under alternative assumptions. Column (4) uses a 5-year window to predict the refusal rate in t based on the average 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 inflated bias-adjusted estimate, although it is not statistically significantly different from the previous estimate. Column (5) adopts a lower value for δ, suggesting that unobservables are relatively less important compared to observables; the resulting estimate is 0.93, which is slightly lower but still close to the baseline IV estimate. Lastly, in column (6), we calculate β_1^* based only on weather and demographic variables as relevant controls, conditioning the estimate on residential planning applications, predicted employment, and local authority and year fixed effects.

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: a majority of homeowners have a strong

TABLE 5 – OWNER-OCCUPIED HOUSING AND HOUSE PRICE APPRECIATION

Dependent variable:	*House price appreciation*						
	Naïve OLS	Baseline IV	+Control variables	+Earnings control	+Demand shock	Homeowners in 1971, 25-75%	Price growth, $t+1$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Majority of homeowners	0.017	0.089***	0.088***	0.054***	0.052***	0.044***	0.060
	(0.012)	(0.025)	(0.029)	(0.013)	(0.009)	(0.012)	(0.038)
Residential planning applications *(log)*	0.001	-0.001	-0.001	-0.001	-0.001	0.000	-0.000
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Male weekly earnings *(log)*				-0.080***			
				(0.021)			
Predicted employment *(log)*					-0.080	-0.104*	0.175
					(0.072)	(0.061)	(0.113)
Weather and demographic controls			✓	✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓	✓
Number of observations	1,416	1,416	1,416	1,416	1,416	1,340	1,770
R^2	0.733						
Kleibergen-Paap F-statistic		8.508	10.28	15.87	12.83	10.87	9.298

Notes: House price appreciation is defined as $\big((\text{House prices}, t+5)/(\text{House prices}, t)\big)/5$ so we calculate the house price growth between $t+5$ and t, excluding data from 2021. In column (7), house price appreciation is based on growth with respect to the next year only, which implies that we do include 2021. Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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 (2)-(7) we instrument for the share of owner-occupied housing with 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. The predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 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. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

impact on regulatory restrictiveness.

3.4 *House price appreciation*

As an important validation exercise of our framework, we also study the effects of homeownership on house price appreciation here. Given that a majority of homeowners imply greater regulatory restrictiveness, we expect that this will push up prices in the area through a limited supply of new housing and amenity effects. We compute the growth in house prices with respect to $t+5$, implying that we have to exclude the last wave of our data (2021). Table 5 reports the results.

The OLS specification does not reveal a strong positive effect of homeownership on price appreciation. However, once we instrument for homeownership in column (2), we observe a robust positive effect: a homeowner majority increases house price appreciation by 8.9% per annum, which is substantial. This effect is likely somewhat overstated, as shown by the coefficient decreasing to 5.4% in column (4) when accounting for local demand. Our preferred specification in column (5), which controls for the demand shock, demonstrates a similar effect of having a majority of homeowners on price appreciation. Column (6) examines LAs with a homeownership share between 25% and 75% in 1971, showing a slightly lower effect of homeowner majority, although not significantly so. Finally, column (7) calculates annual price appreciation, which is inherently noisier than the 5-year measure, but in this way we can include data from 2021. The point estimate aligns with our preferred specifications, but due to larger standard errors, the effect is not statistically significant at conventional levels. Overall, we present strong evidence that a homeowner majority drives substantial price appreciation, likely through increased regulatory restrictiveness.

3.5 Mechanisms

We hypothesized 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 6. 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 homeowner majority on the share of Conservative seats, irrespective of whether homeownership is instrumented for. The preferred specification in column (5) shows that having a majority of homeowners increases the share of Conservative seats by 1.35 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, thereby accentuating their electoral success.

Second, homeowners may seek to directly influence local planning decisions. Efforts may involve filing amendments and petitions, as well as mobilizing neighbours and local communities to protest against specific new developments. Regardless of the outcome of such activities, they inevitably prolong the planning process. To assess this impact, we utilize ancillary data on the share of planning decisions that took longer than 13 weeks. While the OLS specification does not reveal a significant effect, the instrumental variable specifications demonstrate positive and statistically significant effects. The preferred specification in column (11) suggests that a homeowner majority increases the delay rate by almost a standard deviation, indicating a non-negligible impact.

4 A dynamic spatial equilibrium model

4.1 Theoretical framework

4.1.1 Geography

We consider a setting with $i, ..., \mathcal{L}$ discrete locations. Each location is endowed with Λ equally sized plots totalling $\mathcal{L}_i$. Each plot of land is undeveloped, $\Lambda_{it}^{\mathcal{U}}$ or developed $\Lambda_{it}^{\mathcal{D}}$.

Table 6 – Owner-occupied housing, voting and planning delay

Dependent variable:	*Share Conservative council seats (standardized)*						*Residential planning delay rate (standardized)*					
	Naïve OLS	Baseline IV	+Control variables	+Earnings control	+Demand shock	Homeowners in 1971, 25-75%	Naïve OLS	Baseline IV	+Control variables	+Earnings control	+Demand shock	Homeowners in 1971, 25-75%
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Majority of homeowners	0.448***	0.977**	1.035**	1.158**	1.353***	1.130**	0.073	0.674**	0.664**	0.802**	0.976**	1.156**
	(0.095)	(0.488)	(0.523)	(0.491)	(0.505)	(0.443)	(0.057)	(0.262)	(0.262)	(0.338)	(0.435)	(0.547)
Residential planning applications *(log)*	-0.073***	-0.082***	-0.085***	-0.067***	-0.066**	-0.073***	-0.113	-0.116	-0.114	-0.116	-0.115	-0.113
	(0.026)	(0.028)	(0.028)	(0.026)	(0.027)	(0.025)	(0.085)	(0.082)	(0.081)	(0.081)	(0.081)	(0.086)
Male weekly earnings *(log)*				0.946***						0.443**		
				(0.100)						(0.176)		
Predicted employment *(log)*					1.210**	1.326**					0.693	0.822*
					(0.485)	(0.568)					(0.523)	(0.480)
Weather and demographic controls			✓	✓	✓	✓			✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓		✓	✓	✓	✓	✓	
Share council housing 1971 deciles × year fixed effects						✓						✓
Number of observations	1,768	1,768	1,768	1,768	1,768	1,673	1,682	1,682	1,682	1,682	1,682	1,594
R^2	0.763						0.499					
Kleibergen-Paap F-statistic		5.285	5.622	10.18	9.482	7.996		3.556	3.674	6.938	6.470	5.556

Notes: The residential planning delay rate is defined as the share of planning decisions that took more than 8 weeks. Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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 (2)-(6) and (8)-(12) we instrument for the share of owner-occupied housing with 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. The predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 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. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

4.1.2 Production

Production in each location utilise both labour of skill s, ℓ_{it}^{s}, and floor space, f_{it}, to generate output, y_{it}. We assume perfect competition and production operates under constant returns to scale, meaning that output increases proportionally to the inputs used:

$$y_{it}^{s} = \mathcal{A}_{it}^{s}\left(\frac{\ell_{it}^{s}}{\eta}\right)^{\eta}\left(\frac{f_{it}}{1-\eta}\right)^{1-\eta}, \tag{6}$$

where $\mathcal{A}_{it}^{s}$ denotes the productivity in location i at time t. Firms pay wages, w_{it}^{s}, and rents per unit of floor space, p_{it}. Optimal use of inputs implies that:

$$\mathcal{A}_{it}^{s} = (w_{it}^{s})^{\eta}(p_{it})^{1-\eta}. \tag{7}$$

and that the consumption of floor space is given by:

$$f_{it} = \left(\frac{(1-\eta)\mathcal{A}_{it}^{s}}{p_{it}}\right)^{\frac{1}{\eta}}\ell_{it}^{s} \tag{8}$$

4.1.3 Workers

Workers have a specific skill level, s, and choose to be either a homeowner ($\hbar$) or a renter ($\mathfrak{r}$). Their choice of location follows a three-stage decision process. In the first stage, they select a residential location, i. After learning the idiosyncratic shock associated with their chosen residence, they decide on a workplace, j, upon which they then observe a shock specific to the workplace.

Renters. We distinguish between renters and owners. Renters move every year and are myopic. Hence, the flow-utility of renters is given by:

$$u_{it}^{s\mathfrak{r}} = \mathcal{B}_{it}^{s}\left(\frac{f_{it}^{s\mathfrak{r}}}{\theta}\right)^{\theta}\left(\frac{z_{it}^{s\mathfrak{r}}}{1-\theta}\right)^{1-\theta}\xi_{it}^{s\mathfrak{r}}(\omega), \qquad 0<\theta<1, \tag{9}$$

where $\mathcal{B}_{it}^{s}$ represents the amenity level, $f_{it}^{s\mathfrak{r}}$ indicates floor space consumption, and $z_{it}^{s\mathfrak{r}}$ denotes the consumption of the other good. $\xi_{it}^{s\mathfrak{r}}(\omega)$ is an idiosyncratic residential shock that is Fréchet distributed so that $F\big(\xi_{it}^{s\mathfrak{r}}(\omega)\big) = e^{(-\mathcal{B}_{it}^{s}\xi_{it}^{s\mathfrak{r}})^{-\varepsilon}}$. They maximize their utility subject to the budget constraint:

$$\mathbb{E}[w_{it}^{s}] = f_{it}^{s\mathfrak{r}}p_{it} + z_{it}^{s\mathfrak{r}}, \tag{10}$$

where $\mathbb{E}[w_{it}^{s}]$ denotes the expected net wage at i in t and p_{it} are per-period rents.

Given the Fréchet distributed idiosyncratic shock we can write down the share of rents locating at i given the indirect utilities:

$$\pi_{it}^{s\mathfrak{r}} = \frac{(\mathcal{B}_{it}^{s})^{\varepsilon} p_{it}^{-\theta\varepsilon} (\mathbb{E}[w_i^s])^{\varepsilon}}{\sum_{\tilde{i}=1}^{\mathcal{L}} (\mathcal{B}_{\tilde{i}t}^{s\mathfrak{r}})^{\varepsilon} p_{\tilde{i}t}^{-\theta\varepsilon} (\mathbb{E}[w_{\tilde{i}}^s])^{\varepsilon}}. \tag{11}$$

After selecting the residential location, the renter learns about the workplace shock. Consequently, the realised net wage is given by:

$$w_{ijt}^s = e^{-\kappa^s d_{ijt}} w_{jt}^s \xi_{jt}^s(\omega) \tag{12}$$

where $F\big(\xi_{jt}^s(\omega)\big) = e^{(-\mathcal{C}_{jt}^s \xi_{jt}^{s\mathfrak{r}})^{-\varepsilon}}$. This implies that the expected wages are given as:

$$\begin{aligned} \mathbb{E}[w_i^s] &= \frac{\mathcal{C}_{jt}^s (w_{jt}^s)^{\varepsilon}}{\sum_{\tilde{j}=1}^{\mathcal{L}} \mathcal{C}_{\tilde{j}t}^s (w_{\tilde{j}t}^s)^{\varepsilon}} w_j^s \\ &= \pi_{ijt|i}^s w_j \end{aligned} \tag{13}$$

Homeowners. Let us now consider homeowners with skill s who chose to reside in i at time t. Consequently, they maximize utility:

$$u_{it}^{s\mathfrak{h}} = \mathcal{B}_{it}^s \left(\frac{f_{it}^{s\mathfrak{h}}}{\theta}\right)^{\theta} \left(\frac{z_{it}^{s\mathfrak{h}}}{1-\theta}\right)^{1-\theta} \xi_{it}^{s\mathfrak{h}}(\omega), \qquad 0 < \theta < 1, \tag{14}$$

subject to:

$$\mathbb{E}[w_{it}^s] + \mathbb{E}[\mathcal{W}_{it}] = f_{it}^{s\mathfrak{h}} \mu P_{it} + z_{it}^{s\mathfrak{h}}. \tag{15}$$

where μ are mortgage payments and house prices are given by:

$$P_{it} = \sum_{t=0}^{\infty} \frac{p_{it}}{(1+\delta)^t}, \tag{16}$$

so house prices are the present value of the stream of future annual rents. Capital gains with respect to $t+1$ are given by:

$$\mathbb{E}[\mathcal{W}_{it}] = f_{it}^{s\mathfrak{h}} (\mathbb{E}[P_{it+1}] - P_{it}). \tag{17}$$

Combining the budget constraint and expected capital gains lead to the *effective* per-period housing costs:

$$\tilde{p}_{it} = P_{it}(1+\mu) - \mathbb{E}[P_{it+1}]. \tag{18}$$

Then, the optimal consumption of floor space and the composite good are given by:

$$f_{it}^{s\mathfrak{h}} = \frac{\theta \mathbb{E}[w_{it}^s]}{\tilde{p}_{it}} \qquad \text{and} \qquad z_{it}^{s\mathfrak{h}} = (1-\theta)\mathbb{E}[w_{it}^s]. \tag{19}$$

Hence, the expected utility for locating at i is given by:

$$\mathbb{E}[u_{it}^{sh}] = (\mathcal{B}_{it}^{s})^{\varepsilon}\mathbb{E}[w_{it}^{s}]^{\varepsilon}\tilde{p}_{it}^{-\varepsilon\theta}. \tag{20}$$

The homeownership rate at a location i is given by:

$$\begin{aligned} \mathcal{H}_{it} &= \frac{(\mathcal{B}_{it}^{s})^{\varepsilon}\mathbb{E}[w_{it}^{s}]^{\varepsilon}\tilde{p}_{it}^{-\varepsilon\theta}}{(\mathcal{B}_{it}^{s})^{\varepsilon}\mathbb{E}[w_{it}^{s}]^{\varepsilon}\tilde{p}_{it}^{-\varepsilon\theta} + (\mathcal{B}_{it}^{s})^{\varepsilon}\mathbb{E}[w_{it}^{s}]^{\varepsilon}p_{it}^{-\varepsilon\theta}}, \qquad \forall s \\ &= \left(1 + \left(\frac{P_{it}(1+\mu) - \mathbb{E}[P_{it+1}]}{p_{it}}\right)^{\varepsilon\theta}\right)^{-1}. \end{aligned} \tag{21}$$

Hence, when having data on homeownership rates, $\mathcal{H}_{it}^{s}$, rents, r_{it}, and house prices, P_{it}, we can recover expectations about prices with respect to $t+1$, $\mathbb{E}[P_{it+1}]$:

$$\mathbb{E}[P_{it+1}] = P_{it}(1+\mu) - p_{it}\left(\mathcal{H}_{it}^{-1} - 1\right)^{\frac{1}{\varepsilon\theta}} \tag{22}$$

4.1.4 Developers

Consider a land unit, Λ, providing $\bar{f}_i$ units of floor space. Developers decide whether to invest in Λ to create a unit of floor space by maximizing their objective function, $\pi_{it}\bar{f}_i + \epsilon_{it}(\Lambda)$, where $\pi_{it} = (\delta p_{it}\bar{f}_i/d_{it})^{\zeta}$ depends on the rent p_{it} and development costs d_{it}. We assume an idiosyncratic, land unit- and period-specific preference shock, ϵ_{it}, which affects the likelihood of the unit being converted to floor space.

Letting δ denote the discount factor, the value of holding a housing unit in municipality i during period t is given by:

$$\Pi_{it}(\Lambda) = \pi_{it} + \epsilon_{it}(\Lambda) + \delta\mathbb{E}_t\Pi_{it+1}(\Lambda) + \delta\mathbb{E}_t O_{t+1}, \tag{23}$$

were, O_{t+1} denotes the alternative option of housing units located outside the municipality. Consistent with residential sorting models (Bayer et al., 2016; Heblich et al., 2021), we assume that idiosyncratic preferences, ϵ_{it}, follow a Type-I Extreme-value distribution. As a result, the fraction of land developed in i at time t then can be described by a logit model:

$$n_{it} = \frac{e^{\sum_{\tilde{t}=0}^{\infty}\delta^{\tilde{t}}\mathbb{E}_t\pi_{i,t+\tilde{t}}}}{\sum_{j=0}^{\mathcal{L}} e^{\sum_{\tilde{t}=0}^{\infty}\delta^{\tilde{t}}\mathbb{E}_t\pi_{j,t+\tilde{t}}}}. \tag{24}$$

Therefore, the likelihood of a plot being developed depends on expectations regarding future rents and development costs; anticipated increases in development costs will reduce the probability that a plot is developed today.

Please recall that house prices are equal to:

$$P_{it} = \sum_{t=0}^{\infty} \frac{p_{it}}{(1+\delta)^t} \tag{25}$$

Hence, developers are indifferent between remaining owners and renting out the property or selling the property to a homeowner.

We incorporate endogenous development costs that depend on both refusal rates and the current proportion of developed land. This approach is intuitive, as higher regulatory restrictiveness makes planning application outcomes more uncertain, thereby increasing development costs. Additionally, building on previously developed land generally involves greater expenses. Thus,

$$d_{it} = \Phi_{it} + \phi_1 r_{it} + \phi_2 n_{it}, \tag{26}$$

where Φ_{it} is a constant, r_{it} indicates the refusal rate. We expect $\phi_1, \phi_2 > 0$.

4.1.5 Regulatory restrictiveness

The last component of our model addresses the regulatory restrictiveness in each locale i. As demonstrated in Section 3, homeowners exert influence over local planning boards by attempting to sway decisions through voting and by delaying the planning process. Regardless of the specific mechanism, we can represent the refusal rate as a function of whether local homeownership rates exceed the majority threshold. Because it likely takes time before decisions have a measurable effect on regulatory decisions, we expect refusal rates to be affected in the next period.

$$r_{it+1} = \Psi_{it} + \psi_1 \mathbb{I}(\mathcal{H}_{it} > 0.5), \tag{27}$$

where Ψ_{it} is a local constant, $\mathcal{H}_{it}$ is the local homeownership rate (see (**??**)).

The impact of regulation on the spatial economy operates through two primary channels: the supply channel and the amenity channel. The supply channel indicates that increased regulation raises development costs, d_{it} (as shown in (28)), leading to a reduction in the supply of land and floor space. This reduction is expected to drive local house price appreciation, $\mathbb{E}[P_{it+1}]$, which in turn contributes to higher homeownership rates.

The amenity channel suggests that NIMBYism emerges as a means to protect local amenity levels. This is particularly relevant when homeowners value green spaces and are dissuaded by high levels of building density. The amenity channel further implies that:

$$\mathcal{B}^s_{it+1} = \Omega_{it} + \omega_1 r_{it}, \qquad \forall s. \tag{28}$$

which also will lead to elevated levels of local house price appreciation.

4.1.6 General equilibrium

Labour market clearing. Labour market equilibrium requires that the number of in-commuters with a given skill level at workplace location j equals the total workforce at that location by skill level:

$$\sum_{s=1}^{S} \ell^s_{jt} = \sum_{s=1}^{S} \pi^s_{ijt|i} N^s_{it}. \tag{29}$$

Land market clearing. Land market equilibrium requires that the total supply of floor space matches the combined consumption of floor space across tenures and skill levels, in addition to the floor space utilised by the production sector:

$$\begin{aligned} n_i \bar{f}_i &= f_{it} + \sum_{s=1}^{S} f^{s\mathfrak{r}}_{it} + \sum_{s=1}^{S} f^{s\mathfrak{h}}_{it} \\ &= \sum_{s=1}^{S} \left(\frac{(1-\eta)\mathcal{A}^s_{it}}{p_{it}} \right)^{\frac{1}{\eta}} \ell^s_{it} + \left((1-\mathcal{H}_{it}) \frac{\theta \mathbb{E}[w^s_{it}]}{p_{it}} + \mathcal{H}_{it} \frac{\theta \mathbb{E}[w^s_{it}]}{\tilde{p}_{it}} \right) N_{it} \end{aligned} \tag{30}$$

Existence of an equilibrium. For each time period t, the following proposition then holds:

PROPOSITION 1. Assuming strictly positive, finite, and exogenous characteristics, there exist general equilibrium vectors $\{\boldsymbol{R}, \boldsymbol{P}, \mathcal{H}, \boldsymbol{w}^s \ \forall\ s, \boldsymbol{N}, \boldsymbol{M}, \boldsymbol{n}, \boldsymbol{c}, \boldsymbol{r}\}$ for each time period t.

4.2 Estimation and identification

[...]

4.3 Results

[...]

4.4 Counterfactuals

[...]

5 Conclusions

[...]

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Appendix A Theory

Appendix B Data

B.1 *Annual panel dataset*

[...]

B.2 *Additional descriptive statistics*

[...]

TABLE B1 – DESCRIPTIVE STATISTICS FOR THE ANNUAL PANEL DATASET

	Mean	Standard deviation	Min	Max
	(1)	(2)	(3)	(4)
Residential planning refusal rate *(%)*	24.33	17.42	0	100
Residential planning delay rate *(%)*	80.66	19.91	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
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
Mean house price *(£)*	160,032	158,200	15,183	1,982,000
Male weekly earnings *(£)*	468.1	238.4	89.54	1,793
Predicted employment	62,694	51,134	1,663	551,706
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. Those values are available upon request.

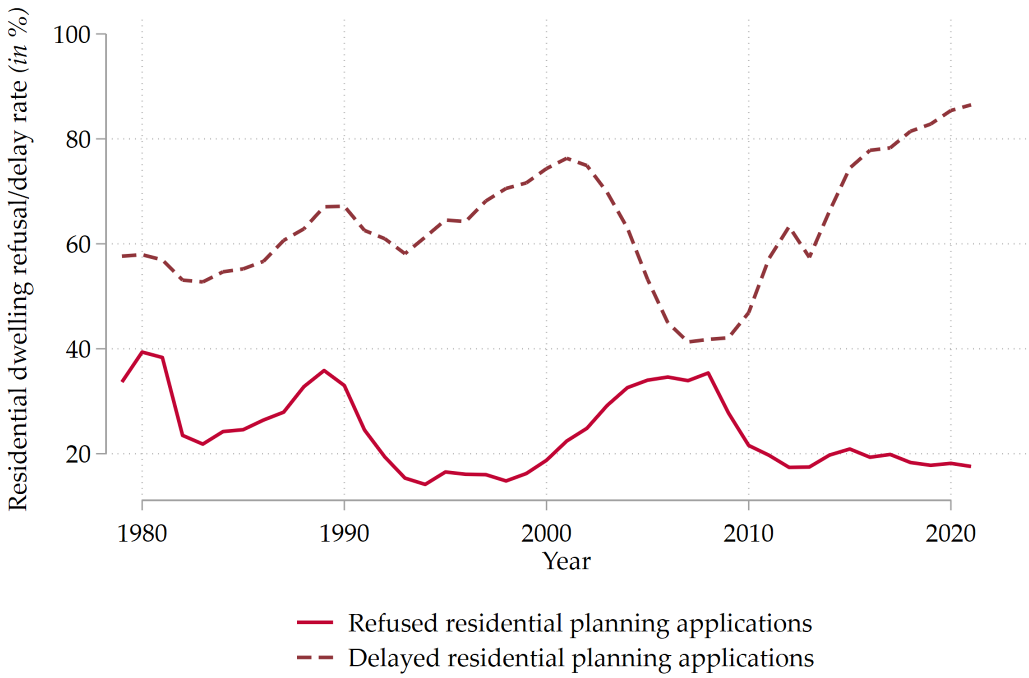


FIGURE B1 – REGULATORY RESTRICTIVENESS OVER TIME

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.

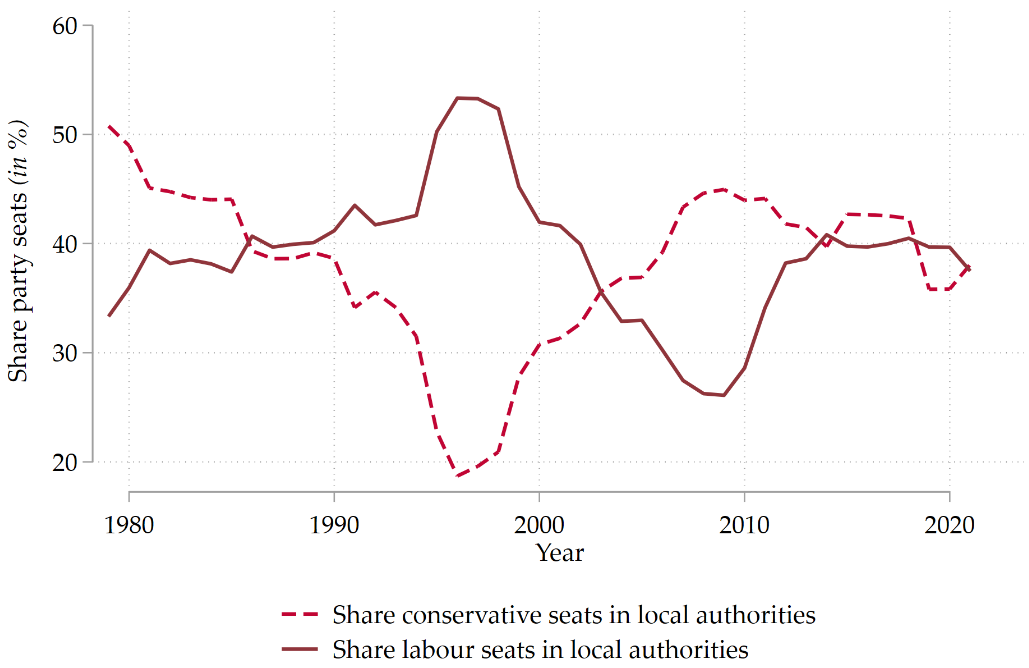
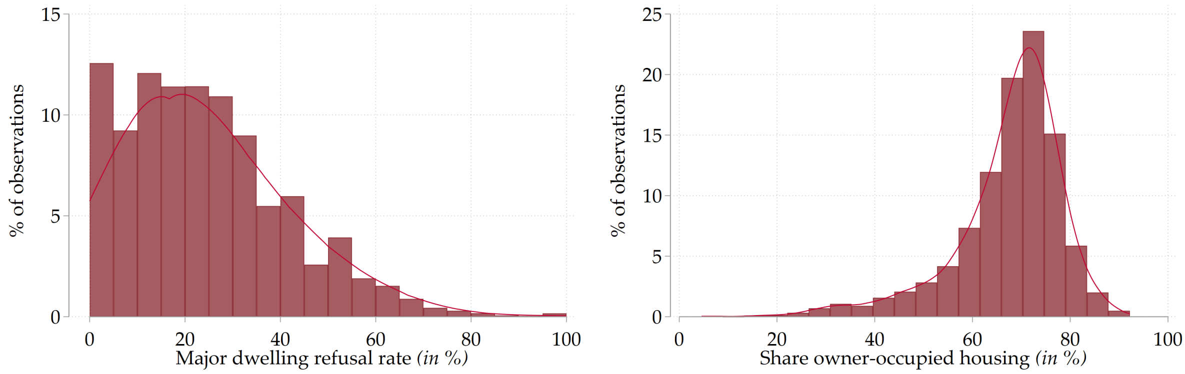


FIGURE B2 – LOCAL ELECTION OUTCOMES BY PARTY

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.

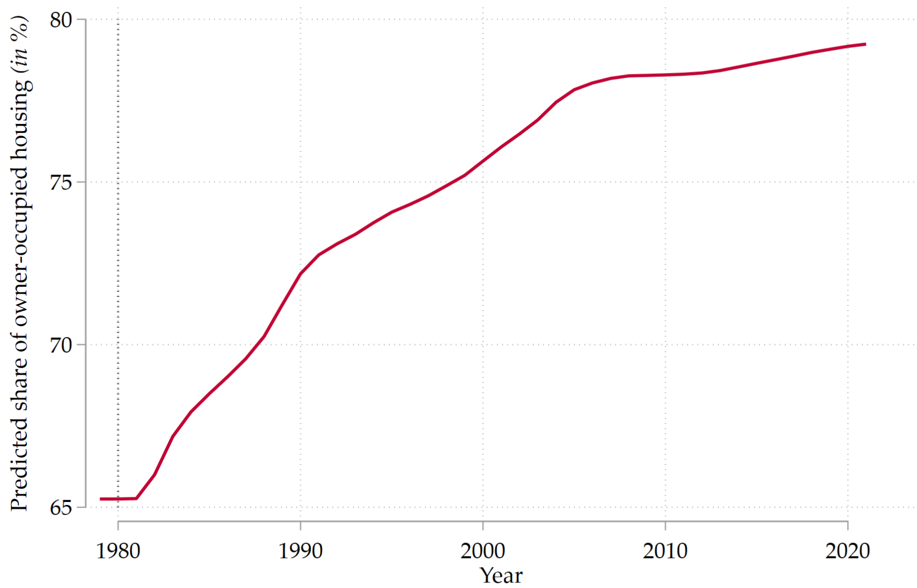
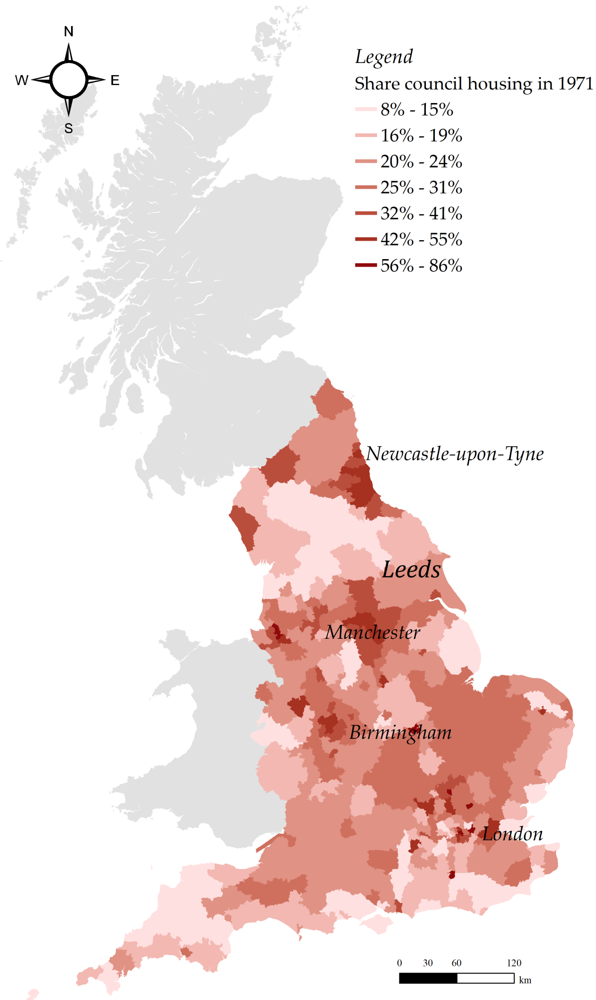
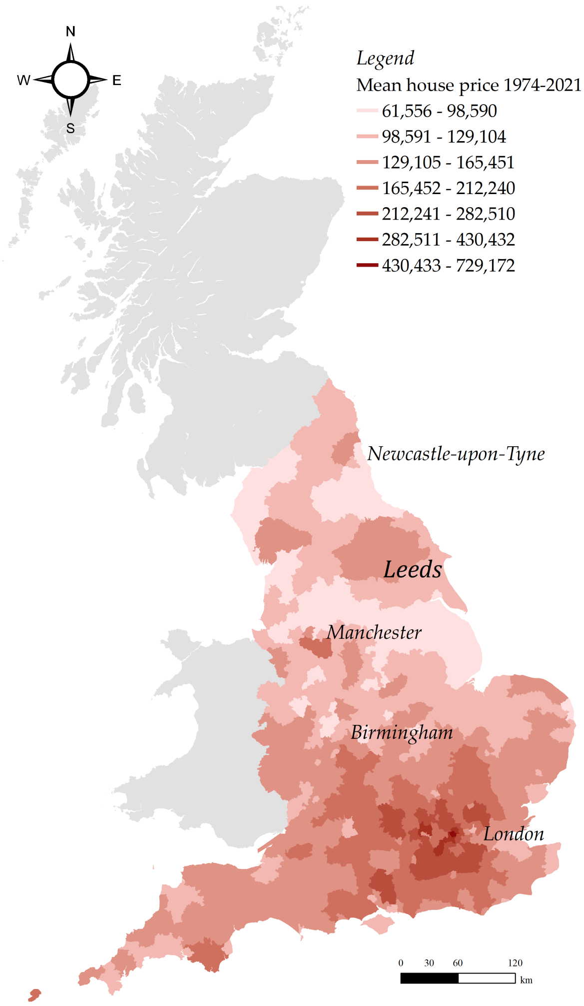


FIGURE B4 – PREDICTED SHARE OF HOMEOWNERSHIP OVER TIME

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 results

C.1 *First-stage results*

TABLE C1 – OWNER-OCCUPIED HOUSING AND REGULATION: FIRST-STAGE RESULTS

Dependent variable:	*Local authority has majority of homeowners*				
	Baseline IV	+Control variables	+Earnings control	+Demand shock	+Homeowners in 1971, 25-75%
	(1)	(2)	(3)	(4)	(5)
Predicted share of owner-occupied housing *(sd)*	0.391**	0.385**	0.380***	0.345***	0.424***
	(0.185)	(0.179)	(0.129)	(0.121)	(0.160)
Residential planning applications *(log)*	-0.000	0.001	0.008	0.007	-0.000
	(0.007)	(0.009)	(0.007)	(0.007)	(0.006)
Male weekly earnings *(log)*			0.025		
			(0.048)		
Predicted employment *(log)*				-0.259***	-0.101***
				(0.065)	(0.039)
Weather and demographic controls		✓	✓	✓	✓
Local authority fixed effects	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓
Share council housing 1971 deciles × year fixed effects					✓
Number of observations	1,682	1,682	1,682	1,682	1,594
R^2	0.721	0.724	0.747	0.748	0.647

Notes: *(sd)* refers to variables standardized with mean zero and unit standard deviation. Weather variables include the minimum temperature, maximum temperature, sunshine hours, and rainfall. Demographic variables include 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 predicted employment is based on 7 industry shares in 1971 in each local authority and national growth rates in these industries since 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. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.