Why Have House Prices Risen So Much More Than Rents in Superstar Cities?*

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Abstract

In most countries house prices have risen much more strongly than rents over the last two decades and much more so in supply-constrained superstar cities. Moreover, the price-to-rent ratio has been cyclical. These facts are consistent with a setting with spatial variation in supply constraints and autocorrelated demand shocks. We test our model predictions employing panel data for England. Our instrumental variable-fixed effect estimates suggest that in Greater London autocorrelated labor demand shocks in conjunction with supply constraints explain two-thirds of the 153% increase in the price-to-rent ratio between 1997 and 2018. We can exclude competing explanations for our findings.

JEL Classification: G12, R11, R21, R31, R52.

Keywords: House prices, housing rents, price-to-rent ratio, price and rent dynamics, housing supply, persistence in demand shocks, land use regulation.

1 Introduction

The new Millennium has brought with it a new crisis: the lack of affordable housing in many urban areas in the developed world, and, particularly in highly productive large cities such as London, New York, Paris, Tokyo, or Hong Kong. The crisis has been profoundly adversely affecting the well-being of residents living in these areas, increasingly causing political unrest locally.

The underlying causes of this affordability crisis, and especially of the strongly rising house prices in so called 'superstar cities' – defined here as desirable (high amenity) cities with severely constrained housing supply growth – have been hotly debated amongst economists, with some pointing to falling real interest rates and others to housing supply shortages. Whether price rises are solely driven by cheaper mortgage financing (so possibly not affecting the affordability of leveraged owner-occupied housing) or by tight land use restrictions and other supply constraints, matters greatly from a policy point of view.

While rising house prices and rents both contribute to the growing affordability crisis, one intriguing stylized fact is that in many – though not in all – countries, house prices have risen much more rapidly over the last two decades than rents. This fact has been employed by some to suggest that there can be 'no supply shortage' as otherwise rents should have risen as much as prices. A second stylized fact is that the increase in the price-to-rent ratio has been cyclical, rising during economic boom periods, but falling during contraction phases.

Figure 1 illustrates these stylized facts of a rising and cyclical price-to-rent ratio for England, France, and the United States (Panels A to C). While in England the house price-to-rent ratio has almost doubled between 1997 and 2018, our sample period, in France and the U.S. it has risen by 84% and 21%, respectively. This stylized fact is even more pronounced for the corresponding superstar cities. In London and Paris, the respective price-to-rent ratios have risen by a staggering 153% and 133%, while in New York City house prices have still grown more than twice as strongly as free-market rents. In all these cases the price-to-rent ratio evolved in a cyclical fashion in line with the business cycle. The dynamics in the price-to-rent ratio is quite different in Japan (Panel D of Figure 1), a country that has been facing an ongoing decline of its population – a sustained negative demand shock. Here the price-to-rent ratio has been falling over the last 20 years, despite a decrease in the real rate of interest. However, in Tokyo, where population has been growing, the price-to-rent ratio increased by 60%.[1]

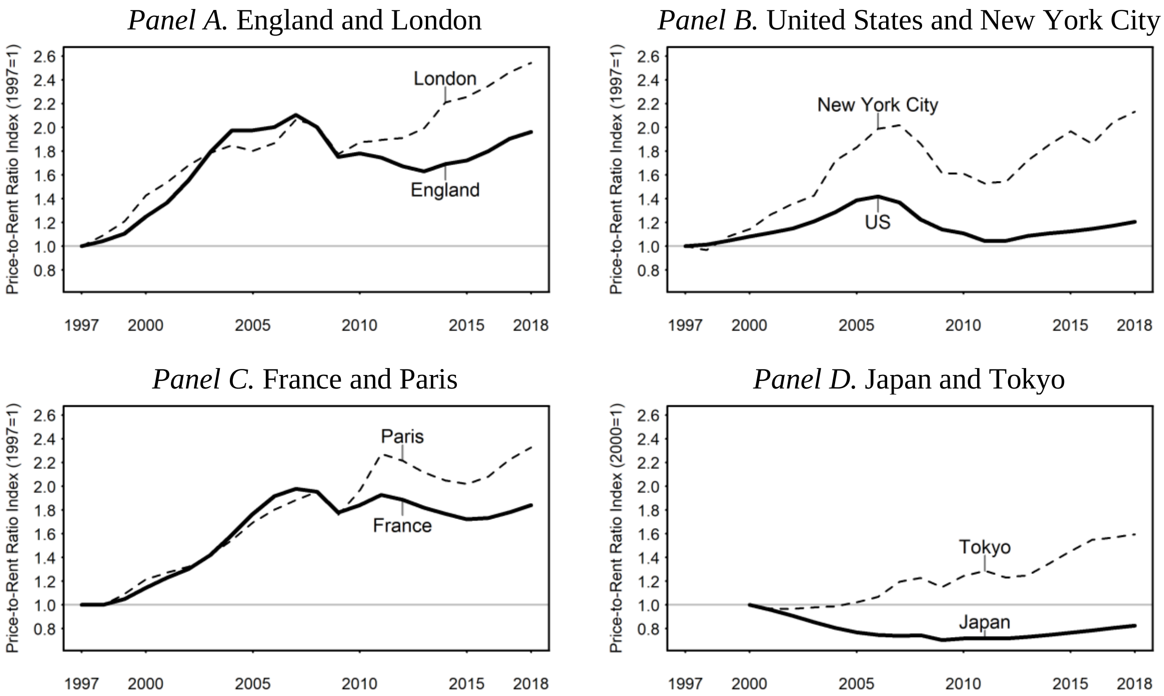
Another fascinating stylized fact is that the price-to-rent ratio since 1997 has been varying enormously across regions within a country. Focusing on the case of England, Figure 2 documents that in Greater London[2] and the South East the increase in the price-to-rent ratio has been far above the country's average, whereas in the North East it has been significantly below.

[1] According to the World Bank, Japan's real interest rate declined from 3.5% in 2000 to 1.1% in 2017. While Japan has a comparably lax planning system, making housing supply fairly price elastic, planning related decisions are more involved in the densely populated metropolitan area of Tokyo. Moreover, lack of developable land severely constrains Tokyo's supply of housing.

[2] When we refer to 'Greater London' or 'London' we mean the Greater London Authority, which consists of 32 local authorities.

Figure 1

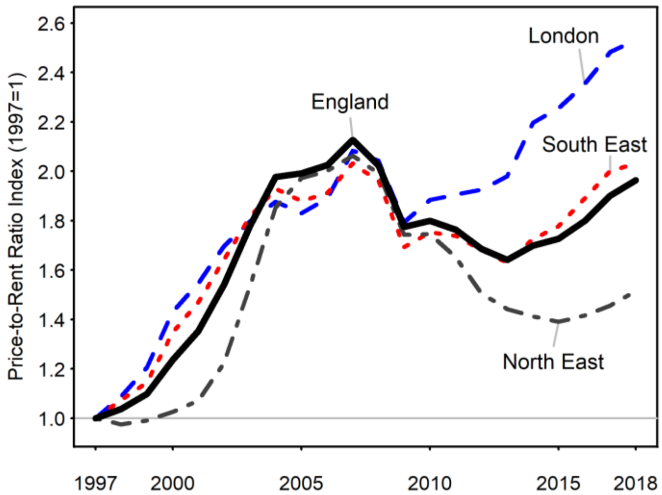
Price-to-Rent Ratio Indices for Selected Countries and Superstar Cities (1997-2018)



Notes: The series for England and London are based on transaction prices (Land Registry) and Private Registered Provider rents (Department for Levelling Up, Housing & Communities, Table 704). The series for the US, France, and Japan are provided by the OECD (data series IDX2015 PRICERENT). The index for New York City is based on the NYU Furman Institute House Price Index for New York City and on a hedonic rent index compiled by the authors, based on mover households in the New York City House Condition and Vacancy Survey (rent controlled units excluded - details are provided on request). For Paris, the rent index is provided by OLAP (free market rents) and the price index is provided by INSEE (transactions of second-hand dwellings, ID 10567012). The city-level index for Tokyo is constructed from hedonic house price and rent indices for the 23 districts of Tokyo (based on Recruit Co. Ltd. listings data; indices provided to the authors by Chihiro Shimizu; see Diewert and Shimizu (2016) for details on the data).

Figure 2

Regional Differences in Price-to-Rent Ratio Indices in England



Notes: The figure displays the ratio of local house prices to rents, averaged over England (black solid line), and the Government Office Regions London, South East, and North East. House prices are based on transactions (Land Registry). Rents are Private Registered Provider rents taken from Department for Levelling Up, Housing & Communities, Live Table 704. We discuss the relationship between market rents and Private Registered Provider rents at length below.

While the unique macroeconomic environment, with a decades-long decline in the real long-term rate of interest or with unprecedented availability of housing credit, could in principle explain a significant increase in the price-to-rent ratio at national level, macroeconomic conditions alone cannot account for the massive differences in the rise of the ratio and its cyclicality at sub-national level.

In this paper we propose a novel theoretical mechanism to explain why house prices can grow more strongly than rents over extended periods of time and why this increase can be expected to be cyclical and much more pronounced in economically thriving and tightly supply-constrained superstar cities, even when holding macroeconomic conditions constant.

Under the standard assumption that the price of a house equals the sum of discounted future expected rental income, the price-to-rent ratio increases either when the discount rate falls or when expected future rental income increases.[3] Our proposed theoretical mechanism focuses on the latter: the determinants of expected future rental incomes.

We develop a simple model where (i) locations differ in their housing supply constraints, (ii) local housing demand shocks exhibit serial correlation – a stylized fact in our data –, and (iii) house buyers have rational expectations about how future expected demand growth affects future rents. In this setting, a positive demand shock increases expected future rent growth, and this increase depends crucially on the location's longer-run housing supply constraints and the degree of serial correlation in the demand shocks: The tighter the supply constraints and the greater the degree of autocorrelation, the greater is the increase in expected future rental income. An increase in expected future rental income, in turn, translates into an increase in the price-to-rent ratio. Moreover, (iv) provided the housing supply curve is inelastic (kinked) 'downwards', the price-to-rent ratio decreases in response to a negative housing demand shock, irrespective of the 'upward' supply price elasticity. All these propositions are consistent with the stylized facts documented in Figures 1 and 2.

In our empirical analysis, we work out the impact of the interaction between local housing demand – proxied in our setting by local labor demand – and local housing supply constraints as well as the dynamics of price- and rent-adjustments during booms and busts in the local economy. To do so, we draw on rich panel data for England over two decades that allow us to study repeated housing booms and busts as well as yearly changes in local labor demand. The latter is an important aspect, as housing demand shocks – driven by labor demand shocks – play a key role in the underlying theoretical mechanism. Moreover, we employ an instrumental variables strategy – building on Hilber and Vermeulen (2016) – to deal with the potential endogeneity of housing supply constraints.

Our empirical focus is on England for three reasons. First, we have extremely detailed data – a unique panel dataset consisting of 353 Local Planning Authorities (LPAs) and annual data from

[3] In the well-known Gordon Growth Model, the price-to-rent ratio depends on the discount rate and the expected long-run future rental growth. The Gordon Growth Model, however, makes important simplifying assumptions (exogenous rent growth, constant rent growth in perpetuity, symmetric response to positive and negative changes in the growth rate) that limit the model's usefulness in explaining the price-to-rent ratio-dynamics in local housing markets, where housing supply elasticities differ. We address these limitations in our theoretical framework.

1974 to 2018 for house prices and from 1997 to 2018 for rents.[4] Second, England provides a particularly relevant laboratory to study the determinants of real house price and rent growth. Since 1970, real house prices have grown more strongly in the UK, and particularly in England, than in any other OECD country and England does not control private rents.[5] Third, partly driven by the severity of the affordability crisis in the most productive and supply-constrained part of the country – Greater London – the political debate of what drives the rising real house prices has been exceptionally fierce.

Our empirical analysis reveals five key insights. First, positive labor demand shocks increase both house prices and rents. The effect on prices is stronger than the effect on rents but factoring in demand shock persistence and the resulting expected future rent increases can account for this difference. Consequently, second, the price-to-rent ratio increases in response to positive labor demand shocks. Third, this effect is much more pronounced in more supply-constrained locations, in line with the proposition that the impact of a demand shock on expected future rental incomes depends crucially on the extent of longer-term supply adjustments. Fourth, the interaction effect with supply constraints is stronger in locations with more persistent demand shocks, consistent with our proposed theoretical mechanism. Fifth, consistent with the supply curve being kinked, *negative* shocks have a negative effect on the price-to-rent ratio, but this negative effect is independent of local supply constraints.

The impact of supply constraints on the price-to-rent ratio is quantitatively important. In Greater London, where supply is seriously constrained, local labor demand shocks in conjunction with local supply constraints explain 64.4% of the increase in the price-to-rent ratio since 1997. The year fixed effects in our panel fixed effects analysis account for the remaining 35.6%. The contribution of local factors vis-à-vis macro factors varies over time. While local supply constraints in conjunction with local demand shocks can explain less than 40% of the increase in the price-to-rent ratio between 1997 and 2008, they can (more than) fully account for the fall in the price-to-rent ratio during the Great Financial Crisis years (2007-2010) and the subsequent increase in the price-to-rent ratio from 2010 to 2018. Consistent with our theoretical propositions, the picture is reversed outside of Greater London, where supply is less tightly constrained. Our simulations suggest that the year fixed effects capture the bulk (84.2%) of the, albeit much smaller, increase in the price-to-rent ratio in the rest of the country.

The year fixed effects are a 'black box'. They are likely to comprise the effect of (i) changing real interest rates and other credit conditions, as well as (ii) the national business cycle in conjunction with aggregate supply constraints. That is, the mechanism we propose applies equally at the aggregate level. Because we standardize the supply constraints-measures in our regressions, the year fixed effects also capture the impact of persistent common housing

[4] LPAs are the local authorities (also called 'councils') that are responsible for the execution of land use planning policy. Given that local regulatory restrictiveness varies across LPAs, they are the logical geographical unit for our analysis. LPAs contain on average 53,158 households, according to the 1991 Census.

[5] Own calculations based on data from the Bank for International Settlement, World Bank and Bank of England. Our analysis focuses on England rather than the entire United Kingdom because consistent planning data over the 45-year horizon is only available for England. Within England, real price growth has been most staggering in London and the South East. London has currently the second dearest buying price of housing per square meter (expressed in US dollars) amongst all prime cities in the world. Only Hong Kong is currently more expensive. See https://www.globalpropertyguide.com/most-expensive-cities, last accessed July 25, 2022.

demand shocks over the business cycle in conjunction with the average housing supply constraints in England. This is especially important in a country like England, where supply constraints are tight by international standards, and where the ‘average location’ arguably has relatively inelastic housing supply. Our empirical model suggests that by 2018, the price-to-rent ratio would have reverted to its 1997 level in a location in England with lax supply constraints, at the first decile of our sample.

One shortcoming of our main data is that we only observe labor demand shocks, but not how these translate into shorter- and longer-term rent growth expectations – a crucial element of our proposed mechanism. To address this limitation and provide insights into the direct link between the two variables, we employ a novel dataset on shorter- and longer-term rent growth expectations of chartered surveyors who advise potential house buyers in England. We document that labor demand shocks affect rent growth expectations more strongly in locations characterized by high demand shock persistence and tight housing supply constraints, consistent with our proposed theoretical mechanism.[6]

We also provide evidence discounting the possibility that our key findings are driven by plausible alternative mechanisms. First, in segmented rental and owner-occupier markets, rising income inequality may increase price-to-rent ratios. Yet, our results are robust to controlling for local income inequality as well as to using several measures of rents and prices that capture the developments in different segments of the housing market, strongly suggesting that local income inequality is not an important factor in our setting. Second, we investigate to what degree changing financing conditions and investment risk may explain our findings. To the extent that changing financing conditions impact the price-to-rent ratio only at the macro-level, we account for them by controlling for year fixed effects in our main specification. However, the impact of financing conditions may vary locally. To address this possibility, we control for location-specific impacts of real mortgage and long-term interest rates by interacting these measures with local supply constraints and land value shares. Our main findings remain virtually unchanged. Similarly, measures of idiosyncratic house price volatility may explain a significant part of the variation in price-to-rent ratios at macro-level. However, location-specific measures of such volatility do not drive the large cyclical swings across locations documented in Figure 2.[7] Third, sticky rents, especially in settings with rent control, may cause prices to increase relative to rents during booms. Yet, there is no rent control in England and regressions in longer time differences yield similar results to our baseline panel fixed effects approach, discounting the concern that our findings may be driven by sticky rents. Finally, the average regression residuals for the local authorities in Greater London are close to zero in each year. This leaves little room for London-specific unobserved factors such as global demand for second homes.

[6] The formation of expectations we observe in the data is consistent with rational behavior. In principle, demand shocks may trigger irrational expectations (an overreaction to highly persistent housing demand shocks in settings with inelastic supply), e.g., during extended boom periods. However, we do not see such patterns in our data.

[7] Changes in real mortgage interest rates and changes in our labor demand measure are positively correlated, suggesting that expansions of credit supply correlating with our labor demand measure are of minor importance. Moreover, our results also hold for the years 2010 to 2018, when the bank rate was stable and real mortgage rates were on a slight positive trend.

The literature on the causes of the growing price-to-rent ratio during the last two decades is scant. The most closely related papers to ours are Van Nieuwerburgh and Weill (2010), Gyourko *et al.* (2013), Büchler *et al.* (2021), Molloy *et al.* (2022), and Kaplan *et al.* (2020).

Van Nieuwerburgh and Weill (2010) develop and calibrate a dynamic equilibrium model of the housing market that features persistence in demand shocks. The model explains well the increase in house price dispersion across metropolitan areas in the U.S. between 1975 and 2007, the main driver being wage dispersion in conjunction with regulatory constraints to local housing supply. Gyourko *et al.* (2013), in their theoretical model, attribute spatial differences in the levels of price-to-rent ratios to spatial differences in long-run expected demand growth in conjunction with relatively inelastic housing supply. Our theoretical mechanism, in contrast, focuses on the substantial cyclical movements in price-to-rent ratios with demand expectations over the local business cycle, and we relate the differences in amplitudes to local supply constraints. Moreover, neither Gyourko *et al.* (2013) nor Van Nieuwerburgh and Weill (2010) focus on rents or the price-to-rent ratio in their empirical analyses.

Büchler *et al.* (2021) and Molloy *et al.* (2022) both explore the role of housing supply elasticities for long-differences in prices and rents during a period of rising housing demand, thereby implicitly ignoring how prices and rents respond to positive and negative demand shocks over the business cycle. Büchler *et al.* (2021) argue that prices react more strongly to demand shocks than rents because shocks lead investors to update their expectations of local risk premiums and rental growth rates, with the degree of updating depending on the share of sophisticated investors at a location. In contrast, we do not rely on exogenous differences in investor beliefs across locations. Our findings are consistent with agents following the *same* rule about updating expectations in all locations. Molloy *et al.* (2022) find that price changes have a stronger association with supply constraints than rent changes, which they explain by households having perfect foresight about long-run demand growth in a setting where future demand growth permanently exceeds the rate of new housing supply. In contrast, in our model agents form rational expectations about the short- and longer-run response to given exogenous shocks, we allow housing supply to eventually catch up to local demand, and we model both positive and negative housing demand shocks.

Kaplan *et al.* (2020), finally, propose a shift in beliefs about future preferences for housing consumption as a driver of price-to-rent ratios during the boom-bust cycle in the U.S. around the Great Financial Crisis. They do not, however, explore why beliefs shifted. We shed light on this: In our setting rational agents form rent growth expectations based on observed demand shocks, local housing supply constraints, and the degree of demand shock persistence. Moreover, this channel links the differences in local housing supply constraints to the cyclical spatial differences in price-to-rent ratios that we observe in the data.

Our paper also ties into – and helps reconcile disagreements between – different strands of a growing literature on the root causes of the housing affordability crisis that has emerged since the late 1990s, especially in superstar cities. Broadly speaking there are two main propositions.

The first strand, largely an urban economics literature, highlights the supply side and the micro-location; in particular, the role of binding local land use restrictions. It suggests that the rise in real house prices in desirable cities is largely the result of tighter local planning constraints in conjunction with strong positive demand shocks in these locations. Most studies focus on the United States and find a causal effect of land use regulation on house prices (e.g., Glaeser and Gyourko 2003, Glaeser *et al.* 2005a and 2005b, Quigley and Raphael 2005, Glaeser *et al.* 2008, Saks 2008), in particular, in desirable larger cities, so called ‘superstar cities’ (Gyourko *et al.* 2013). In the UK, various reviews and studies (e.g., Cheshire and Sheppard 2002, OECD 2004, Barker 2004 and 2006) suggest that the decades-long undersupply of housing and the extraordinary growth in real house prices is linked to a dysfunctional planning system. Hilber and Vermeulen (2016) provide rigorous empirical evidence for England suggesting a causal effect of local regulatory constraints on the real house price-earnings elasticity. Other related work (Cheshire and Hilber 2008) points to the tax system and the lack of tax-induced incentives at the local level to permit development.

The second strand emphasizes the demand side and the financing of housing, with a focus on the United States. It argues that a unique macroeconomic environment with a decline in the real rate of interest, unprecedented availability of housing credit, and/or global investor demand for superstar locations may jointly explain much, if not all, of the increase in real prices.[8] Himmelberg *et al.* (2005) suggest that it was easily available credit in the years preceding the Great Financial Crisis that boosted housing demand and house prices. Favara and Imbs (2015) demonstrate that branching deregulations in the US between 1994 and 2005 led to positive credit supply shocks driving up house prices, and more so in areas with inelastic housing supply. In a similar vein, Justiniano *et al.* (2019) provide stylized facts for the boom years that are consistent with looser lending constraints (shifts in credit supply), but not with looser borrowing constraints (shifts in credit demand). Greenwald and Guren (2020) suggest that changing credit conditions may explain a significant fraction of the increase in the price-to-rent ratio over the boom years. Overall, this literature provides persuasive evidence that credit supply plays an important role in explaining the house price boom in the US prior to the Great Financial Crisis.

In the UK, deregulation of credit markets occurred much earlier than in the US. In fact, the most significant changes relating to housing credit ensued before the start of our sample period, between 1983 and 1997. Arguably, the most important reform step was the Finance Act in 1983, which abolished the interest rate cartel of so-called ‘building societies’. Deregulation therefore does not appear to be the driver of the growth in real house prices and of the price-to-rent ratio in England since 1997.[9] Similarly, the cyclical patterns we observe are hard to rationalize with sustained improvements in financing conditions over time due to innovation. Moreover, our findings also hold for the sub-periods from 2007 to 2009 and from 2010 to 2018 – that is, during

[8] Other studies however question the importance of falling real interest rates in explaining the house price boom preceding the Great Financial Crisis. Apart from Kaplan *et al.* (2020), discussed above, Favilukis *et al.* (2017) suggest it was the relaxation of financing constraints (generated entirely through a decline in the housing risk premium) rather than lower interest rates that led to the boom. Glaeser *et al.* (2012) document that lower real interest rates can explain only one-fifth of the rise in US house prices between 1996 and 2006.

[9] Recent work in the UK has instead focused on the persistent decline in real interest rates over the last two decades and the tightening of credit conditions in 2008. Miles and Monro (2021) argue that the surge in aggregate house prices in the UK between 1985 and 2018 was driven by increasing incomes and an unexpected fall in the real interest rate, with both components being equally important, but they do not consider the role of housing supply constraints for the cyclical differences in price-to-rent ratios across locations.

and after the Great Financial Crisis. In particular, from 2007 to 2009 real mortgage interest rates and long-term interest rates both fell, at a time when the price-to-rent ratio decreased notably.

The core contributions of our study to these strands of the literature are threefold. First, our study reconciles the urban economics and macro/finance strands of the literature by proposing and testing a novel theoretical mechanism that is consistent with both growing real house prices and rents, and growing price-to-rent ratios during the past two decades, especially in supply-constrained superstar cities like London. Our study stresses the importance of local demand and supply side determinants, alongside macroeconomic factors captured in our empirical setting by year fixed effects. Second, and related, we provide evidence that long-run supply constraints are quantitatively important in explaining rising price-to-rent ratios during extended housing boom periods and we discount the possibility that this is driven by alternative mechanisms. This also refutes the narrative of some proponents in the policy arena who interpret the rising price-to-rent ratio as 'direct evidence' that the housing shortage is not driven by a dysfunctional planning system or lack of developable land.[10] Third, we shed light on the cyclical nature of the price-to-rent ratio, stressing the importance of the persistence of demand shocks in explaining strongly growing price-to-rent ratios in economic expansion periods as well as falling ratios during busts.

Our paper is structured as follows. In Section 2, we present our theoretical model and formulate propositions. Section 3 discusses the underlying data and our identification strategy. We then present results of our baseline specifications, explore alternative explanations, and conduct robustness checks. In Section 4, we investigate the quantitative importance of the mechanism. The final section concludes.

2 Theory

To explain why not only house prices and rents but also the price-to-rent ratio responds more strongly to labor demand shocks when housing supply is tightly constrained, we develop a simple stylized model of local housing markets that differ in their short- and long-run housing supply elasticities. The purpose of this model is to illustrate the theoretical mechanism we have in mind. This mechanism builds on three crucial assumptions. First, local housing demand shocks exhibit serial correlation (Assumption 1), which is a feature of our data. Second, we assume that locations differ in their longer-run housing supply constraints and that locations with tighter long-run constraints also face comparably tighter short-run supply constraints than locations with laxer long-run constraints (Assumption 2). This is because of binding short-run planning and construction lags.[11] Third, we assume that agents in the market form rational expectations based on these relationships (Assumption 3).

Since contemporaneous market rents only depend on short-term demand and supply, the slope of the short-run supply curve will determine the effect of a housing demand shock on

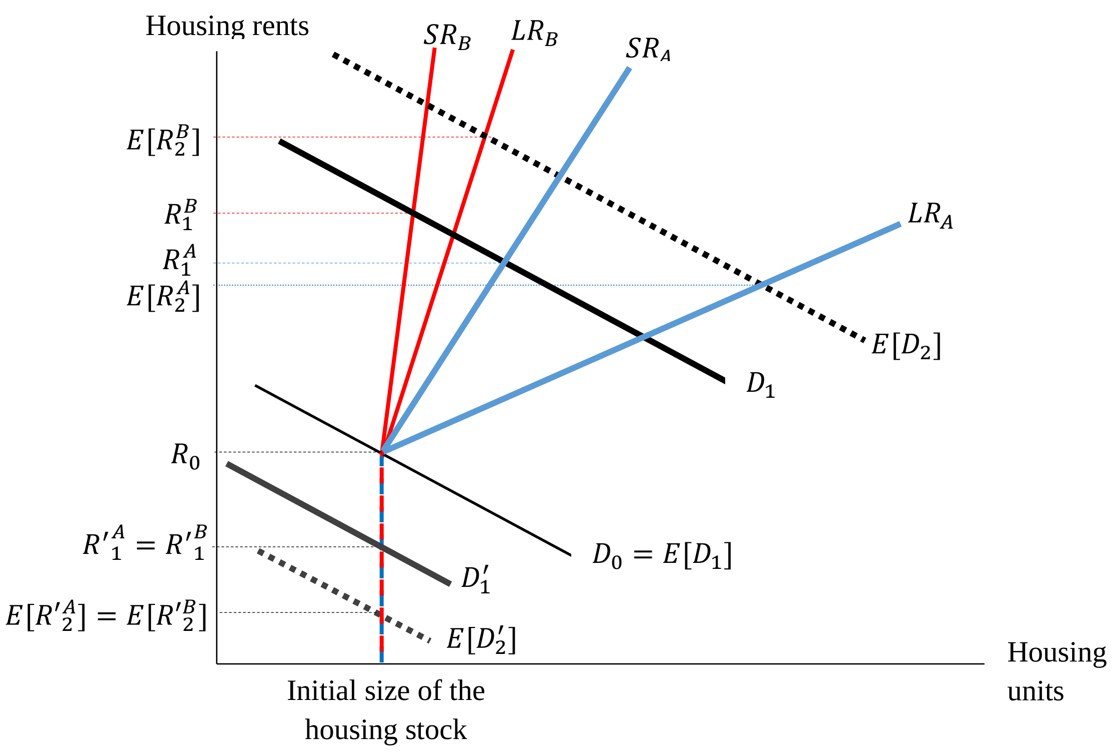
[10] The most prominent proponent in England is Mulheirn (2019). See also Been *et al.* (2019) who critically assess the 'supply skepticism' arguments in the United States.

[11] There are several reasons for this: First, the delay rate of planning applications increases with general regulatory restrictiveness. Second, it is harder for developers to find adequate open land for development if a location is already more built-up, and construction takes longer if the developer must tear down an old building before being able to start development. Third, it is more difficult to build in locations that are more rugged, which arguably

contemporaneous rents. Markets with less elastic short-run housing supply will experience a stronger rent increase in reaction to a given positive demand shock than markets where housing supply is more elastic in the short-run. Absent of demand shocks being serially correlated, the rent level will be higher in the short- than the long-run. This is because the new housing supply triggered by the demand shock shifts the new market equilibrium to the right eventually. However, with positive serial correlation (Assumption 1) and depending on the elasticity of housing supply, future expected rents may be higher despite the larger long-run supply elasticity. In that case, prices react more strongly to the initial demand shock than rents.

Figure 3 illustrates this point. In location A, the upper parts of the housing supply schedules for short- and long-run housing supply are less steep (more elastic) than in location B. The lower parts are vertical in both locations, representing the durability of housing (Glaeser and Gyourko 2005). A positive demand shock in period 1 (the short run), shifting the demand schedule from D_0 to D_1, increases rents up to the intersections with the respective short-run supply curves. Since supply is more elastic in location A, rents increase less sharply there. Due to the serial correlation of the demand shock, the expected long-run demand, $E[D_2]$, is to the right of the short-run demand curve. The intersections of $E[D_2]$ with the long-run supply curves, LR_A and LR_B, determine the expected long-run rent levels. As long as the autocorrelation of the demand shock is sufficiently strong to outweigh the attenuating effect of the long-run supply expansion, rents are expected to increase further. In the example depicted in Figure 3, this is the case in location B, but not in A.

Figure 3
Theoretical Mechanism



increases construction time. For all these reasons, short- and long-run elasticities are highly likely positively correlated.

In equilibrium in period 0, E$[R_1]$ = R_0, hence P_0 = $R_0(1 + 1/(1 + r))$, where r is the discount rate and P_0 is the price that an investor would be willing to pay for the house. The demand shock shifts the price to P_1=R_1 + E$[R_2]/(1 + r)$ in period 1. The price-to-rent ratio increases in response to the demand shock if $P_0/ R_0 < P_1/ R_1$, that is if R_1< E$[R_2]$. Consequently, the price-to-rent ratio increases in location B, but falls in A. The underlying reason is the difference in the supply price elasticity. In contrast to a positive initial demand shock, a negative demand shock, D'_1, has the same quantitative impact in both locations because of the kink in the housing supply schedule, implying an equal decrease in the price-to-rent ratio in both locations.

We now turn to the model. We start with a simple setting where the housing supply schedule does not exhibit a kink. In this case, the reaction to a negative shock can be expected to be a mirror image of the reaction to a positive shock. We then discuss the case of a kinked supply curve (as depicted in Figure 3), where the housing supply elasticity is zero below the equilibrium point. This alters the prediction for negative shocks.

2.1 Model Economy: Case of Symmetric Housing Supply Schedule

We consider a representative location in a modified Rosen (1979) and Roback (1982) framework. The model has an initial period and two main periods. In the initial period 0, the location is in spatial equilibrium, so that the location's population is constant. In that situation, its wage rate is hit by a shock. We then consider the short-run reaction of housing demand and supply to the shock (period 1 denotes the short-run; 1 year), before discussing the (expected) longer-run equilibrium outcome (period 2 captures the longer-run; 5-10 years). This setting with only two main periods has the advantage of being simple while still maintaining the key mechanism we have in mind.[12]

Assume that the location is characterized by a short- and a long-run housing supply elasticity, which we take to be exogenous[13], as well as by location fundamentals a_t (amenities) and ω_t (wages). We define $w_t = a_t + \omega_t$ as the amenity-adjusted wage rate in period t. The location's initial housing stock is S_0. We assume that households in the model are renters. Investor-landlords willing to pay the present discounted value of the housing unit determine the price of housing.[14]

Households are indexed by i. They have an outside option that yields utility $\bar{u}$, which we normalize to $\bar{u} = 0$, and they are mobile in all periods.

[12] In a setting with an infinite number of periods, the key results from our simple setting could be maintained in numerical simulations if one were to impose a per-period construction capacity limit, as in Wheaton (1999).

[13] The short- and long-run supply price elasticities are determined by geographical, topographical, and regulatory constraints. In our empirical work we deal with the endogeneity of these determinants by employing an IV-strategy.

[14] By assuming that the price is determined by investor-landlords, we ignore the possibility that rental and owner-occupied markets may be perfectly segmented. This is a potential concern because we empirically observe the price paid by owner-occupiers. If the rise in the price-to-rent ratio were driven by increasing incomes for owner-occupiers but stagnating or falling incomes for renters, this too could explain an increase in the price-to-rent ratio over our sample period. Empirically, we show in Section 3.5 that changes in local income inequality do not alter our main findings. Theoretically, in a strict sense, markets are only segmented if renter and owner-occupier households never switch between markets. Switchers (e.g., first-time homebuyers) contribute to arbitrage between

The locations in our corresponding empirical analysis are small relative to the country. Consistent with this, we ignore the impact of shocks to the representative location on the utility associated with the outside option (other locations) in the model. This is a common assumption in the urban economics literature.

Household i's utility from living in the location in period t is $w_t - R_t - \eta_i$, whereby R_t is the rent and households have an idiosyncratic (dis-)taste for the location, represented by $\eta_i \sim \mathcal{U}_{[0,\phi]}$. If ϕ is small, households have a relatively stronger taste for the location, on average. We assume that ϕ is large enough so that some households prefer to live elsewhere.

Households with draws $\eta_i \leq \bar{\eta}$ choose to live in the representative location, so that housing demand is given by

$$D_t = \int_0^{\bar{\eta}} \frac{1}{\phi} d\eta = \frac{\bar{\eta}}{\phi} = \frac{1}{\phi}(w_t - R_t). \quad (1)$$

The resulting initial equilibrium rent level in period 0 is $R_0 = w - \phi S_0$. The location is in equilibrium, defined as a situation where current housing demand and supply, and expected future housing demand are equal.

In period 1, the location's wage rate is hit by a shock ε that also entails information about the evolution of wages in period 2. The expected change in the wage rate in period 2 is given by $\gamma\varepsilon$, where $\gamma \in (-1,1)$ captures the degree of autocorrelation of the demand shock.

Housing developers can react to the shock in period 1 by expanding housing supply. The (reduced-form) short-run housing supply function is given by

$$S_1 = S_0 + \delta\beta(R_1 - R_0). \quad (2)$$

Following Mayer and Somerville (2000), this supply function captures the idea that housing developers react to price changes, rather than the level of prices. The parameter $\delta \in (0,1]$ governs the difference between short- and long-run housing supply. A smaller δ means that short-run supply is less elastic relative to long-run supply of the location, whereas $\delta = 1$ represents the case where the full supply response occurs in the short-run and therefore the short-run and long-run supply curves are identical. β captures the location's long-run supply elasticity. A smaller β reduces both the short- and the long-run supply elasticity. This merely implies that, if the short-run supply curve is more elastic in one location than the other, the same is true for the long-run supply curve (Assumption 2).

Equating short-run supply and demand $D_1(\varepsilon)$, and solving for the equilibrium rent yields

$$R_1 = R_0 + \frac{1}{1+\phi\delta\beta}\varepsilon. \quad (3)$$

This expression shows that rents increase in response to a positive demand shock ($\varepsilon > 0$), and this increase is more pronounced if local short-run housing supply is less elastic (i.e., when $\delta\beta$ is small), and if demand is more elastic (i.e., when ϕ is small).

the two segments, helping to equalize housing cost differentials. The mechanism we propose also applies if market segmentation exists but is imperfect.

After having observed the demand shock ε, the long-run expected demand is $E[D_2] = (w + \varepsilon(1 + \gamma) - E[R_2])/\phi$, which follows from the linearity of the demand curve. $E[R_2]$ is the expected long-run rent level. The long-run supply curve is $S_2 = S_0 + \beta(E[R_2] - R_0)$. This yields an expected long-run rent level

$$E[R_2] = R_0 + \frac{1+\gamma}{1+\phi\beta}\varepsilon. \tag{4}$$

The long-run expected rent also increases in response to a positive demand shock ($\varepsilon > 0$), and more so if local housing supply is less elastic and local housing demand is more elastic relative to other locations (i.e., when β or ϕ are small). Therefore, similar relationships hold for the house price conditional on having observed the shock in period 1, which we define as $P_1 = R_1 + E[R_2]/(1 + r)$. We summarize these predictions in propositions.

PROPOSITION 1. *Consider a positive housing demand shock, ε>0. House prices increase and the increase is more pronounced if housing supply in the location is relatively inelastic compared to other locations. (If the housing supply schedule is symmetric around the equilibrium point, an analogous statement applies in the case ε<0.)*

PROPOSITION 2. *Consider a positive housing demand shock, ε>0. Housing rents increase, and the increase is more pronounced if housing supply in the location is relatively inelastic as compared to other locations. (If the housing supply schedule is symmetric around the equilibrium point, an analogous statement applies in the case ε<0.)*

2.2 Price-to-Rent Ratio: Case of Symmetric Housing Supply Schedule

In the initial situation, the price-to-rent ratio is simply $1 + 1/(1 + r)$. The price-to-rent ratio increases in response to a positive local housing demand shock if $E[R_2] > R_1$, which is the case for

$$\gamma > \frac{\phi\beta(1-\delta)}{1+\phi\beta\delta} \in (0,1). \tag{5}$$

That is, if the housing demand shock is sufficiently strongly autocorrelated, the expected increase in future demand outweighs the long-run supply response. This is more likely if β is small, which reduces the long-run supply response, or if local housing demand is relatively elastic (i.e., when ϕ is small).[15] In that case, the autocorrelated housing demand shock will have a relatively stronger impact on future housing demand.

Finally, the impact of the housing demand shock on the price-to-rent ratio becomes more positive when housing supply is more inelastic. Stronger autocorrelation of the demand shock amplifies this mechanism. We summarize these results as follows:

[15] The housing demand price elasticity across English regions has been shown to be rather uniform around -0.4 to -0.5 (see Ermisch *et al.* 1996).

PROPOSITION 3. *Consider a small positive housing demand shock, ε>0.*

(i) The price-to-rent ratio increases in response to a positive demand shock if demand shocks are sufficiently strongly autocorrelated.

(ii) The impact of the housing demand shock on the price-to-rent ratio becomes more positive when housing supply is more inelastic.

(iii) Stronger autocorrelation of the demand shock amplifies the interaction effect of the demand shock with the housing supply elasticity.

(If the housing supply schedule is symmetric around the equilibrium point, an analogous statement applies in the case ε<0.)

Proof: See Appendix A.

The proposition includes the case where the full supply response happens in the short-run, i.e., when $\delta = 1$. Then, we would expect the price-to-rent ratio to always increase in response to a positive demand shock. For $\delta < 1$ the prediction is more subtle: The price-to-rent ratio may decrease in response to a positive demand shock in locations where either persistence is low, or supply is elastic. As Sections 3 and 4 show, this latter case is relevant empirically, even in parts of the overall comparably tightly supply-constrained England.

2.3 Price-to-Rent Ratio: Case of Kinked Housing Supply Curve

For ease of exposition, the above discussion focused on positive labor demand shocks. This would be sufficient if the housing supply schedule were symmetric around the equilibrium point. However, there are good reasons to believe that, because of the durability of the housing stock, supply is much less elastic when housing demand shocks are negative (Glaeser and Gyourko 2005). We refer to this setting as a 'kinked supply curve'.

Consider a negative shock to housing demand, $\varepsilon < 0$. If the supply cure is vertical below the equilibrium point in all locations, we have $S_1 = E[S_2] = S_0$, so that $(w_t - R_t)/\phi = S_0$ for $t = 1, 2$. Hence, $R_1 = w + \varepsilon - \phi S_0$ and $E[R_2] = w + \varepsilon(1 + \gamma) - \phi S_0$, which shows that prices and rents decrease in response to a negative housing demand shock. The price-to-rent ratio decreases if $E[R_2] < R_1 \Leftrightarrow \varepsilon\gamma < 0$. This is true as long as the labor demand shocks exhibit positive serial correlation, i.e., $\gamma > 0$.

PROPOSITION 4. *Suppose that the housing supply schedule has a kink at the equilibrium point. Consider a situation with a negative initial housing demand shock, ε<0.*

(i) House prices decrease. The decrease is independent of the upward supply price elasticity of the location.

(ii) Rents decrease. The decrease is independent of the upward supply price elasticity of the location.

(iii) The price-to-rent ratio decreases (as long as the housing demand shock exhibits positive serial correlation). The decrease is independent of the upward supply price elasticity of the location.

The simple two-period model does not address the question whether the predicted effects are permanent or transitory. Standard asset pricing theory implies that they are transitory.

Nonetheless, the example of the sustained decline in population in Japan – causing the price-to-rent ratio to fall – suggests that these transitions can last for decades.[16]

According to *Proposition 3*, differential positive rent growth expectations emerge during local housing booms because the persistence of demand shocks gives rise to further growth expectations, and because supply constraints amplify the impact of the current and expected positive demand shocks on current and future rents, respectively. As Figures 1 and 2 show, these local upswings can last a decade (and possibly longer). Rent growth expectations revert when the boom ends, removing transitory spatial differences in price-to-rent ratios. *Proposition 4* suggests that price-to-rent ratios move in tandem in all locations during the bust – again consistent with Figures 1 and 2.

3 Empirical Analysis

3.1 Data and Descriptive Statistics

We compile a panel data set at LPA-level covering the years 1974 to 2018 for house prices and 1997 to 2018 for rents. Summary statistics of the key variables are reported in Table 1. We provide more detail and background information in Online Appendix O-A.

House Prices, Rents, and Price-to-Rent Ratio

The main outcome variable in our analysis is the price-to-rent ratio at LPA-level. We construct this variable from housing transaction prices and rents, deflating all nominal values by the national-level retail price index net of mortgage payments (RPIX). For the house price series, we build on Hilber and Vermeulen (2016) and use transaction data from the Council of Mortgage Lenders (1974-1995) and the Land Registry (1995-2018) to calculate mix-adjusted real house price indices at LPA-level. We refine the index by dropping 'Right to Buy' transactions[17] from the Council of Mortgage Lenders data. The full house price series covers the period from 1974 to 2018.

We employ two measures for local rents. The first is the mean private market rent, provided by the Valuation Office Agency. Private market rents are available from 2010 to 2018. We construct a mix-adjusted index that holds constant the average dwelling size (number of rooms).

While private market rents are in principle our preferred measure of local rents, the fact that this data is only available for nine years is a significant limitation. We therefore employ a second measure – Private Registered Provider (PRP) rents – that yields a much longer time-series. PRP rents are provided by the Department for Levelling Up, Housing & Communities (DLUHC) and are available from 1997 to 2018.[18] While some PRPs are for-profit organizations, others are not-for-profit. In all cases however, they have an incentive to maximize their rental income, subject to constraints; not-for-profit organizations to be able to reinvest surplus income

[16] Our model considers local business cycles and local supply constraints. However, the mechanism of our model in principle also applies to the macro-level (e.g., for Japan as a whole). Put differently, what we capture with the year fixed effects in our empirical model could either be driven by interest rates or by our mechanism at the macro-level (or, in fact, by some other macroeconomic factor, such as changing credit conditions).

[17] The 'Right to Buy' scheme, implemented in 1980, permitted tenants in Council Housing to buy their homes at a discount that could be as high as 40% of the market value of the unit.

[18] 1997 is the first year with any available rental data for England at local level. See the gov.uk Live Table 704.

to provide additional housing. All PRPs face a rent ceiling. This ceiling is typically defined as a fraction of the market rent that a particular unit would obtain on the free market.

Table 1

Summary Statistics

	Mean	Standard Deviation			Min.	Max.
		Overall	Between	Within		
A. LPA panel, 1974-2018 (N = 353, T = 45)						
Mix-adj. real house price index (1974 = 100) [a)]	194.2	97.3	29.1	92.8	23.7	1015.7
Log(local labor demand) [b)]	10.76	0.65	0.64	0.07	8.15	13.16
Help to Buy (post-2015) x London dummy	0.006	0.079	0.019	0.076	0	1
B. LPA panel, 1997-2018 (N = 353, T = 22)						
Mix-adj. real house price index (1974 = 100) [a)]	268.6	85.8	53.3	67.2	99.2	1015.7
Real weekly rents (PRP rents in £)	96.1	14.7	12.8	7.2	58.9	151.4
Ratio of house prices to yearly PRP rents	50.7	22.5	19.3	11.5	15.2	327.5
Ratio of dwelling prices to yearly PRP rents	33.1	16.9	14.4	9.0	5.9	243.9
Log(local labor demand) [b)]	10.8	0.64	0.64	0.05	8.35	13.16
Idiosyncratic house price risk (Giacoletti 2021)	0.258	0.056	0.044	0.034	0.085	0.546
Idiosyncratic house price risk (own measure)	0.138	0.030	0.024	0.018	0.071	0.513
C. LPA panel, 1997-2018, harmonized/outliers removed (N = 344, T = 22)						
Mix-adj. real house price index (1974 = 100) [a)]	246.3	76.3	41.9	63.7	99.2	1015.7
Real weekly rents (PRP rents in £)	95.6	14.3	12.5	6.9	58.9	151.2
Ratio of house prices to yearly PRP rents	48.6	16.6	13.0	10.4	15.2	126.5
Ratio of dwelling prices to yearly PRP rents	31.5	11.5	8.3	8.0	5.9	103.7
Log(local labor demand) [b)]	10.8	0.62	0.62	0.05	9.26	13.16
Idiosyncratic house price risk (Giacoletti 2021)	0.258	0.056	0.044	0.035	0.085	0.546
Idiosyncratic house price risk (own measure)	0.137	0.028	0.023	0.017	0.071	0.336
D. LPA cross-section (N = 353)						
Avg. refusal rate of major resident. projects, 1979-2018	0.242	0.083			0	0.473
Share of greenbelt land in 1973	0.088	0.215			0	1
Change in delay rate b/w 1994–96 & 2004–06	-0.031	0.220			-0.635	0.531
Share of votes for Labour, 1983 General Election	0.163	0.091			0.001	0.410
Share of developable land developed in 1990	0.257	0.233			0.009	0.976
Population density in 1911 (persons per km²)	733.3	2562			3.250	2.2e5
Range between highest and lowest altitude (m)	208.8	171.2			5.000	975.0
Land value share [c)]	0.114	0.77			0.029	0.982
E. Government Office Regions panel, 2013-2018 (N = 9, T=6)						
One-year ahead expected real rent growth	-0.004	0.015	0.006	0.014	-0.041	0.032
Five-years ahead expected real rent growth	0.014	0.017	0.006	0.016	-0.018	0.051

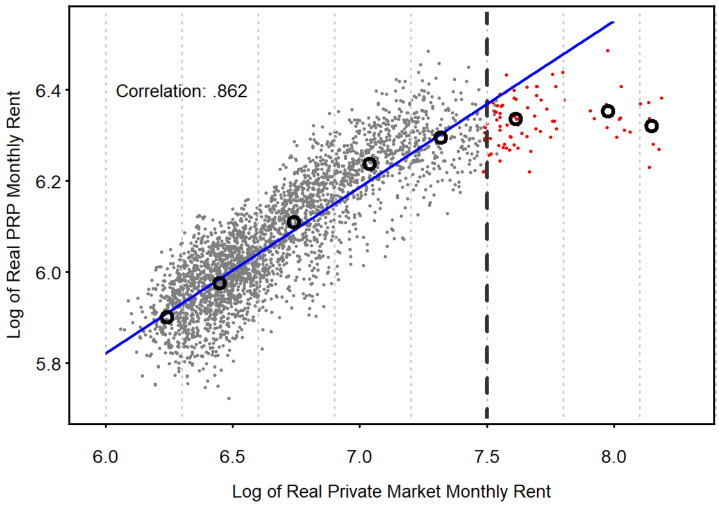
Notes: [a)] Based on house price transaction data. [b)] Log predicted employment, based on 1971 local industry composition and national employment growth. [c)] Approximated using data on land values and house prices.

Because PRP rents allow us to cover a period of 22 years, with several (local) booms and busts, as opposed to only 9 years, we employ this measure in our baseline analysis.

One may be concerned that PRP rents are not a good proxy for market rents. To assess this, Figure 4 depicts a scatterplot of the two measures by LPA and year, suggesting a strong positive relationship, except for LPAs with a very high private market rent (to the right of the dashed vertical line). The figure also displays averages for equally sized bins (bold black rings) that further support this conjecture. This suggests that PRP rents adequately capture the private market rent dynamics for most of the LPAs in our sample.

Figure 4

Private Registered Provider and Market Rents Scatterplot and Correlation



Notes: The graph plots the log of the real market monthly rent against the log of the real Private Rental Provider monthly rent, by LPA and year. The bold black rings represent averages for the bins defined by the vertical light grey dashed bars. Each bin has a width of .3, starting at 6.0. The red dots indicate LPAs excluded from the regression sample because the relationship between the two types of rents seems to differ from the relationship in other LPAs. Average log real market rents in these LPAs exceeded 7.5.

To deal with the possibility that PRP rents may not adequately proxy for private market rents in LPAs with very high private market rents, we use a simple exclusion-rule based on a visual inspection of Figure 4: We drop all LPAs with a mean log market rent exceeding 7.5.

In our baseline regression, we thus measure the local price-to-rent ratio as the ratio of average house prices to average PRP rents. As Table 1 shows, this measure of the price-to-rent ratio produces relatively large values, a direct consequence of the PRP rent ceiling. We run a series of robustness checks using alternative rent and price measures: (i) private market rents only, (ii) PRP rents derived from different sample selection rules based on the correlation between changes in PRP rents and market rents, (iii) dwelling prices, and (iv) a repeated-sales price index. These regressions show that our baseline results are robust to the choice of the rent and price measures and are not sensitive to leaving out the high-rent LPAs to the right of the vertical dashed grey line in Figure 4.

Housing Supply Constraints

We use three measures as proxies for the long-run supply price elasticity. Building on the literature (Burchfield *et al.* 2006, Saiz 2010, Hilber and Vermeulen 2016) we employ measures that capture regulatory, physical/geographical, and topographical long-run supply constraints, respectively. Our measure of regulatory restrictiveness is the average refusal rate of major residential planning applications from 1979 to 2018 derived from the DLUHC. The 'refusal rate' is simply the number of refused 'major applications' (i.e., applications of projects consisting of ten or more dwellings) divided by the total number of such applications in a given year. This is the standard measure used in the literature to capture regulatory restrictiveness in Britain – see Hilber and Vermeulen (2016). Our two other supply constraint-measures are taken from Hilber and Vermeulen (2016): the share of developable land already developed in 1990 and the range in elevation in the LPA, as a proxy for terrain ruggedness. Steep terrain and ruggedness make building costlier, and thus represent a physical constraint to housing supply. The refusal rate and share developed measures are arguably endogenous. We discuss our instrumental variable strategy to identify the causal effects of these two measures below.

Measure of Local Housing Demand

Our proxy for local housing demand is a Bartik (1991) shift-share measure that captures local labor demand. For LPA i and year t, local labor demand is given by $LLD_{it} = \sum_{k=1}^{7} emp_{i,1971}^{k} \times index_t^k$, where $emp_{i,1971}^{k}$ is employment in industry k in 1971, and $index_t^k$ is the national-level employment index for industry k (base year 1971). Following Hilber and Vermeulen (2016), we take employment by industry for seven industries at LPA-level from the 1971 Census and national-level employment growth by industry from the Census of Employment (1971-1978) and the Office for National Statistics (1979-2018).

Our theoretical model assumes that shocks to local housing demand exhibit persistence. To test this assumption in the data, we first regress the change in the log labor demand on the lagged change in the log labor demand and a constant, separately for each LPA and based on the full period from 1974 to 2018. Figure 5 summarizes the spatial distribution of the autocorrelation parameter. The variation across LPAs is not particularly large, with 79% of LPAs exhibiting autocorrelation between 0.5 and 0.6. Moreover, all LPAs in London fall into this range, as indicated by the red vertical bars, suggesting that our main finding (London stands out) may not be driven by a different level of persistence in the demand shock in the capital.[19]

3.2 Endogeneity Concerns and Identification Strategy

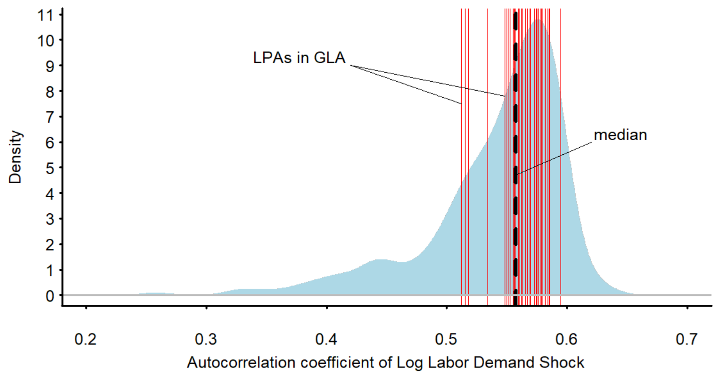
To capture the mechanism proposed by the theoretical model, we need to isolate exogenous variation in local housing supply constraints from local housing demand and other confounders. Our strategy to identify the causal effects of local supply constraints is three-pronged.

First, we exploit the panel structure of our data: We control for time-invariant confounders through location (LPA) fixed effects, and we capture the impact of common macroeconomic shocks through year fixed effects.

[19] The median degree of persistence hardly differs between locations inside and outside of the Greater London area (0.566 and 0.556).

Figure 5

Spatial Distribution of Persistence in Local Labor Demand Shocks



Notes: The graph plot the spatial distribution of autocorrelation coefficients of the local labor demand shock. Vertical red bars indicate LPAs in London. The dashed vertical black line is the median of the full sample, 0.558. Autocorrelation coefficients were computed by LPA, using the entire labor demand series covering 1974-2018.

Second, the shift-share measure of local housing demand (i.e., our predicted local labor demand measure) transforms time-series variation at the national level into local shocks that are arguably orthogonal to the state of the local housing market. As noted above, our baseline period is 1971, pre-dating our sample period by over a quarter of a century. One advantage of our shift-share measure compared to using local earnings as demand shifter is that it cannot be influenced by house prices through income sorting and therefore it may only reflect housing demand and not housing supply. One concern with it is that the initial industry composition in a location may correlate with unobserved shocks to the relative attractiveness of renting versus owning. Another concern is that the financial sector is an important driver of local labor demand shocks in some LPAs and that the shift-share measure thus may capture local credit availability as well. We deal with these threats to identification in the robustness check section.

Third, we use an instrumental variable strategy to identify the causal effects of local housing supply constraints. One general threat to the identification of supply constraints is that they tend to be correlated with housing demand conditions (Davidoff 2016). Other endogeneity concerns relate more specifically to our measures of regulatory restrictiveness and scarcity of developable land. We discuss how we deal with these concerns below.

Identifying Regulatory Supply Constraints

Our measure of local regulatory restrictiveness is the average share of planning applications for major residential projects that are refused by the elected councilors in an LPA over the period from 1979 (the first year with available data) to 2018. Our implicit assumption is that LPAs that tend to refuse a higher share of projects, are more restrictive in nature (rather than that they are faced with consistently poorer planning applications).

We follow Hilber and Vermeulen (2016) and use the *average* local refusal rate from 1979 to 2018, instead of annual data. We do so for two reasons. First, refusal rates are highly pro-cyclical. All else equal, higher demand for housing should lead to a higher number of planning applications. However, the capacity of LPAs to process applications is likely limited. From the perspective of the LPA, one strategy to deal with the excess workload could be to reject some applications quickly. We would thus expect to see a greater share of rejections during boom periods and indeed this is what the data conveys. Second, a developer wishing to build in a very restrictive LPA likely faces higher (expected) administrative costs of applying and a lower chance of approval. If a developer feels that the chances of a rejection are high, she might spend more time working out applications for projects that have a fair chance of acceptance and submit a smaller total number of applications in the first place. In this case, the refusal rate underestimates the true regulatory restrictiveness.

We may still be concerned however that even the average refusal rate is endogenous. After all, planning decisions are the outcome of a political economical process (Hilber and Robert-Nicoud 2013). We thus employ three quite different instruments and demonstrate that our results are robust to changing the combination of instruments used.

Our first instrument is the LPA share of greenbelt land in 1973, 24 years prior to the start of our sample period for the price-to-rent ratio analysis.[20] Greenbelt land is de facto protected from development, but it constitutes a large share of the land around many English cities. For instance, Greater London covers 157k hectares in total, of which around 35k hectares are greenbelt land. While this is already a substantial share, the whole London Metropolitan Greenbelt covers 514k hectares of land, more than four times the non-greenbelt area of Greater London. The situation is similar in other English cities, such as Liverpool and Manchester. Clearly, this represents a major obstacle to new development. LPAs that were assigned a large share of greenbelt land back in 1973 arguably were also those with strong cohorts of Not-in-My-Backyard (NIMBY)-residents who would subsequently fight hard to maintain the status quo. Thus, we may expect that the share of historic greenbelt land and subsequent restrictive local planning are strongly positively correlated. However, the historic share of greenbelt land should not directly affect contemporaneous changes in the price-to-rent ratio (other than through regulatory restrictiveness). The facts that (i) this instrument substantially predates the sample period and (ii) greenbelt land is used for agricultural rather than recreational purposes,[21] makes it unlikely that contemporaneous changes in demand conditions that correlate with the refusal rate also correlate with the instrument.

Our instruments two and three were initially proposed by Hilber and Vermeulen (2016). The second instrument stems from a reform of the English planning system in 2002 that created exogenous variation in local regulatory restrictiveness. The reform imposed a speed-of-decision target for major developments onto LPAs. Prior to the reform, a more restrictive LPA could simply delay the decision instead of rejecting an application; delays and rejections were effectively substitutes. The reform sanctioned delays, but planning authorities were still allowed to reject applications.[22] Hilber and Vermeulen (2016) show in their figure 1 that prior to the reform, changes in the refusal rate and changes in the delay rate were uncorrelated, that is, all planning parameters were optimized in pre-reform equilibrium. The reform then prompted a temporary strong negative correlation between the change in the delay rate and the change in the refusal rate before eventually the two measures became uncorrelated again. This implies

[20] We calculate the share of (protected) greenbelt land in 1973 from a digitized historic map of Great Britain (Lawrence 1973) and a shapefile of the 2001 LPA boundaries. See Online Appendix O-A for more information.
[21] Greenbelts may not be confused with public parks, which are the main recreational attractions in English cities.
[22] The sanctions were implicit rather than explicit, see Hilber and Vermeulen (2016).

that restrictive LPAs – to meet their delay rate target – responded to the reform by delaying less and refusing more.[23]

Our identifying assumption is that the reform had a differential impact on more and less restrictive LPAs: The most restrictive LPAs should have had the strongest incentive pre-reform (between 1994 and 1996) to delay residential applications and the strongest incentive post-reform (between 2004 and 2006) to reduce their delay rate and instead refuse more applications. While the refusal rate is endogenous, our instrument – the change in the delay rate (post- vs. pre-reform) – is a policy-induced exogenous source of variation in regulatory restrictiveness. Our instrument 'change in delay rate' can be expected to be strongly correlated with the endogenous average refusal rate (measured between 1979 and 2018), yet we would not expect the change in the delay rate to directly – other than through regulatory restrictiveness – affect contemporaneous changes in the price-to-rent ratio.

Our third instrument is the vote share of the Labour party in the 1983 General Election (derived from the British Election Studies Information System). This and similar instruments have been used previously to identify planning restrictiveness (Bertrand and Kramarz 2002, Sadun 2015, Hilber and Vermeulen 2016). On average, voters of the Labour party have below-average incomes and housing wealth and they are more likely to rent. We would thus expect this group to care less about the protection of housing wealth, and more about the affordability of housing. This suggests a negative correlation between the Labour vote share and local planning restrictiveness, all else equal. Our identifying assumption is that the share of Labour votes affects house price and rent changes only through its impact on local restrictiveness, after controlling for LPA and year fixed effects. By using general election results, pre-dating the sample period of our main analysis by 14 years, we minimize the threat that local demand conditions or development projects at the local level influence the election results. Hence, outcomes of the planning process most likely did not determine the election outcomes that we use as instrument.

In our baseline regression, we use the three instruments jointly. In robustness checks, we explore the sensitivity of the results to using only two or one of the three instruments.

Identifying the Share of Developed Land

The share of developable land developed in 1990 captures the degree to which new development is likely to be costly redevelopment rather than more straightforward development on greenfield land. The measure is potentially endogenous to local demand conditions. Some places may have become more attractive over time because of better amenities or economic opportunities, leading to immigration from less desirable locations. This would result in a higher share of developed land in 1990. Likewise, the planning decisions of an LPA prior to 1990 may influence the amount of open land in 1990. To deal with these potential sources of endogeneity, we adopt the strategy proposed by Hilber and Vermeulen (2016) and instrument the share of developed land in 1990 with population density in 1911. The rationale is that population density in 1911 is indicative of (time-constant) local amenities and the productivity of a place (which predicts the share of developed land almost 80 years later), but the effect of

[23] LPA-level delay rates are published by the DLUHC.

this on average house prices and rents in an LPA will be captured by the LPA-fixed effects. On the other hand, we do not expect historic population density to be correlated with changes in contemporaneous demand conditions. It is thus unlikely that historic density influences changes in house prices and rents during our sample period through other channels than scarcity of land.

3.3 Empirical Baseline IV-Specification

The theoretical model developed in Section 2 suggests that the impact of local housing demand shocks on local house prices, rents, and the price-to-rent ratio depends on local housing supply constraints. We estimate the following fixed effects specification:

$$\begin{aligned} y_{it} &= \theta_0 \log LLD_{it} + \theta_1 \log LLD_{it} \times \overline{refusal\,rate}_i + \theta_2 \log LLD_{it} \times \%developed_i \\ &\quad + \theta_3 \log LLD_{it} \times elevation_i + HTB[i \in London] \times I(t > 2015) + LPA_i + year_t + e_{it}. \quad (7) \end{aligned}$$

We include LPA and year fixed effects in all regressions, to control for time-constant local differences in housing-related variables as well as macroeconomic factors that vary over time, but not locally. As outcomes y_{it}, we consider a log mix-adjusted real house price index, log real rents, and the price-to-rent ratio for LPA *i* and year *t*.

The main source of variation comes from our measure of local housing demand, the natural logarithm of predicted local labor demand, LLD_{it} (i.e., our shift-share measure). Although this variable enters in levels, since we control for LPA fixed effects, it has the same interpretation as a first difference specification and hence captures shocks to local labor demand.

To allow for a differential impact of local demand shocks on the outcomes, we interact $\log LLD_{it}$ with the average refusal rate of major residential projects in LPA i, $\overline{refusal\,rate}_i$, the share of developable land already developed in 1990, $\%developed_i$, and the elevation range, $elevation_i$.

All three measures enter in standardized form (i.e., normalized to the mean being equal to zero and the standard deviation being equal to one), so that the interpretation of the coefficients $\theta_0, \ldots, \theta_3$ is straightforward: θ_0 captures the impact of a labor demand shock on the outcome in an LPA with average supply constraints in all three dimensions. The coefficients θ_1, θ_2, and θ_3 capture the additional impact of a local labor demand shock when the respective supply constraint increases by one standard deviation.

We instrument for the interaction of the refusal rate by the interactions of the labor demand shock with the three instrumental variables for the refusal rate (the share of historic greenbelt land, the reform-based change in the delay rate, and the share of Labour votes in the 1983 General Election). The instrument for the share developed land is the historic population density in 1911.

The regressions also control for a dummy $HTB[i \in London] \times I(t > 2015)$ that is equal to one for LPAs in London observed after 2015. The dummy captures the differential impact of a recent housing market policy in England: Help to Buy. Introduced in England in 2013, the policy aims to help households to purchase a home, with the main instrument being an equity loan scheme. From 2016 onwards, the policy was more generous in London, relative to the rest of the country (Carozzi *et al.* 2020).

We estimate this main specification for the baseline sample as well as for periods with positive and negative housing demand shocks, respectively.

3.4 Main Results

Prices and Rents

Before turning to the price-to-rent ratio as outcome variable, we consider the impact of local supply constraints and labor demand shocks on real house prices and rents separately. Table 2 displays our baseline results, testing *Propositions 1* and *2*. The dependent variable in column (1) is the log mix-adjusted real house price index and estimation is by OLS. This ignores endogeneity concerns related to the local regulatory restrictiveness and the share developed land measures. The period covered is the full sample period for the house price data, 1974-2018. The log LLD as well as the interaction terms with the refusal rate and the share developed land are highly significant and positive (consistent with *Proposition 1*), and so is the Help to Buy dummy. The altitude range interaction is insignificant and close to zero.

Table 2

Impact of Labor Demand Shocks on House Prices and Rents

	(1)	(2)	(3)	(4)
	Log(Prices) OLS 1974-2018	Log(Prices) 2SLS [a] 1974-2018	Log(Prices) 2SLS [a] 1997-2018 [b), c)]	Log(Rents) 2SLS [a] 1997-2018 [c]
Log(local labor demand, LLD)	0.556***	0.317**	-0.067	0.022
	(0.092)	(0.132)	(0.155)	(0.129)
Av. refusal rate of major residential projects × log(LLD)	0.188***	0.652***	0.863***	0.283***
	(0.069)	(0.118)	(0.123)	(0.071)
Share of developable land developed in 1990 × log(LLD)	0.438***	1.099***	1.110***	0.504***
	(0.148)	(0.117)	(0.253)	(0.083)
Range between highest and lowest altitude × log(LLD)	-0.044	0.326***	0.203*	0.124**
	(0.041)	(0.108)	(0.122)	(0.056)
Help to Buy (post-2015) x London dummy	0.242***	0.047*	0.035	-0.049***
	(0.065)	(0.027)	(0.046)	(0.015)
LPA FEs	Yes	Yes	Yes	Yes
Year Fes	Yes	Yes	Yes	Yes
Observations	15,885	15,885	7,555	7,555
Number of LPAs	353	353	344	344
R-sq. overall	0.027			
R-sq. within	0.960			
R-sq. between	0.138			
Kleibergen-Paap F		17.89	9.747	9.747

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. [a] First stage results are reported in Table B1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). [b] Observations with missing rental data removed to make price and rent specifications comparable. [c] PRP vs. market rent outliers (mean log market rent > 7.5, based on Figure 4) removed.

In column (2), we estimate the same regression by Two-Stage Least Squares (2SLS), instrumenting the refusal rate- and share developed-log LLD interactions. In this regression, the independent effect of log LLD and its interaction with the supply constraints are positive and highly significant. Moreover, the supply constraint interactions are quantitatively more important, as compared to the OLS-estimates. As noted above, if a developer expects LPAs to reject a project, the developer might consider not to apply for planning approval in the first place. This would lead to an underestimation of the true refusal rate and could be one of the reasons why the coefficient on the interaction term in the OLS specification is lower. The Kleibergen-Paap F statistic does not show signs of weak instruments, and the coefficients are very similar to those obtained by Hilber & Vermeulen (2016). This is despite extending the sample by ten years, using a refined house price series that accounts for discounted transactions under the Right-to-Buy scheme, and adding the share greenbelt instrument for improved identification of regulatory restrictiveness.

We report the corresponding first stage regression results in columns (1) and (2) of Table B1. (All subsequent first stage results corresponding to Table 2 are also reported in Table B1.) In all first stage regressions, the share of greenbelt land in 1973, the reform-based change in the delay rate, and the Labour party vote share correlate strongly and in expected ways with the refusal rate of major residential projects. Similarly, the historic population density in 1911 is a strong predictor of the share of developable land already developed in 1990.

In column (3) of Table 2, we repeat the regression in column (2) for the sub-period and LPAs covered by the rental data. The interaction terms do not change much, but the main effect of the labor demand measure turns slightly negative and becomes insignificant.

In column (4), the outcome variable is the log real PRP rents. Here, we restrict the sample to LPAs where the average log market rent 2010-2018 does not exceed 7.5 (see Figure 4).[24] Qualitatively, the results look very similar to the price regression results (consistent with *Proposition 2*), but all interaction terms are smaller in magnitude. This suggests that local housing supply constraints play a relatively larger role in shaping the impact of local labor demand shocks on house prices, as suggested by the theoretical model (presuming that the local labor demand shocks are sufficiently strongly autocorrelated).

To make sense of the relative differences between column (3) – the impact of a demand shock on prices – and column (4) – its impact on rents – we provide back-of-the-envelope calculations in Table 3 that factor in the magnifying effect of demand shock persistence. Column (1) corresponds to the case when the degree of persistence is at the sample median of 0.558. In Panel A, we assume a discount rate of 2% consisting of a 1.5% risk premium and a risk-free rate of 0.5%, to capture the low-interest rate environment after the Great Financial Crisis. In this scenario, taking the share of developable land as an example, the initial impact of the demand shock on rents plus the expected future rent increases due to demand shock persistence imply a price coefficient of 1.112 – very close to the actual price coefficient of 1.110 (also reported in column (4) of Table 3 for easier comparison).[25] The cumulative impact changes slightly when shifting the degree of persistence by one standard deviation below or above the median in columns (2) and (3).

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[24] As discussed below, we conduct several robustness checks that use a more refined approach. We also show results for the full sample, using market rents. We use the 7.5 log points threshold in our baseline analysis because it is straightforward. However, the results do not hinge on this choice.

[25] $\sum_{t=0}^{\infty} 0.504 * 0.558^t * (1/1.02)^t = 1.112$. If we assumed a 5-year horizon instead of infinity, the implied coefficient would be very similar (1.083).

Table 3

Price Coefficients Implied by Rent Regression (Column (4) of Table 2) and Demand Shock Persistence

Degree of persistence	Price coefficient implied by rent regression			Actual price coefficient (95% CI in brackets)
	Median (1)	–1 SD (2)	+1 SD (3)	(4)
Panel A. Scenario 1 – Discount rate = 2%				
Av. refusal rate × log(LLD)	0.624	0.556	0.711	0.863 [0.622, 1.104]
Share developable × log(LLD)	1.112	0.991	1.266	1.110 [0.614, 1.606]
Altitude range × log(LLD)	0.273	0.244	0.311	0.203 [-0.036, 0.442]
Panel B. Scenario 2 - Discount rate = 6.5%				
Av. refusal rate × log(LLD)	0.594	0.534	0.668	0.863 [0.622, 1.104]
Share developable × log(LLD)	1.058	0.952	1.190	1.110 [0.614, 1.606]
Altitude range × log(LLD)	0.260	0.234	0.293	0.203 [-0.036, 0.442]
Panel C. Scenario 3 - Discount rate = 6.5% and attenuating supply response				
Av. refusal rate × log(LLD)	0.521	0.559	0.486	0.863 [0.622, 1.104]
Share developable × log(LLD)	0.928	0.996	0.865	1.110 [0.614, 1.606]
Altitude range × log(LLD)	0.228	0.245	0.213	0.203 [-0.036, 0.442]

Notes: Columns (1) — (3) show the cumulative impact of expected rent changes following an initial labor demand shock, based on the coefficients from column (4) of Table 2. The degree of persistence in column (1) is 0.558, the sample median over all LPAs. The other columns consider persistence one standard deviation (0.056) below and above the median, to capture the role of statistical uncertainty in the measurement of persistence. For easier comparison, column (4) displays the actual price coefficients from column (3) of Table 2 with 95% confidence intervals. In Panel A, we assume a discount rate of 2%, corresponding to the low-interest rate environment after the Great Financial Crisis (assuming a bank rate of 0.5% and a risk premium of 1.5%). In Panel B, we use a discount rate of 6.5%, consisting of the median bank rate 1997—2007 of 5% a risk premium of 1.5%. In Panel C, we assume a discount rate of 6.5% and reduce expected future rental price changes linearly from 100% in year 0 to zero in year 10, to capture the attenuating effect of the supply response.

In Panel B, we repeat this exercise with a higher discount rate of 6.5%, motivated by the average bank rate of 5% between 1997 and 2007 and a risk premium of 1.5%. This does not markedly affect the cumulative impact because the demand shock persistence decays quickly.

The first two scenarios ignore the fact that long-run supply is likely to expand more than short-run supply in response to the demand change. In Panel C, we incorporate the attenuating effect of future supply adjustments by reducing the expected future rent increases from 100% in year 0 to 0% in year 10, in steps of ten percentage points, using the 6.5% discount rate. Even in this 'conservative' third scenario, the price coefficients implied by the rent regression are reasonably close to the estimated price coefficients. Overall, Table 3 reveals that expected future rent changes in conjunction with demand shock persistence can explain the differential impact of the demand shock on rents and prices rather well, in line with our theory. Moreover, the fairly closely matching coefficients leave little scope for irrational exuberance.

Price-to-Rent Ratios (Baseline Estimates)

In a next step we consider the price-to-rent ratio as the outcome variable in Table 4, testing *Proposition 3 (ii)* by regressing the price-to-rent ratio on the same set of explanatory variables as in Table 2. The results reveal that the price-to-rent ratio increases in an average LPA in response to a positive local labor demand shock and the impact of this shock is stronger when regulatory (refusal rate) and physical (share developed land, altitude range) supply constraints are tightened, consistent with the proposition.

Table 4

Determinants of Price-to-Rent Ratio (Baseline Specification)

	Price-to-rent ratio 2SLS [a)] 1997-2018 [b)] Baseline
Log(local labor demand)	39.441*** (10.617)
Av. refusal rate × log(local labor demand)	60.149*** (8.805)
Share developed × log(local labor demand)	79.275*** (17.921)
Altitude range × log(local labor demand)	22.755*** (8.529)
Help to Buy (post-2015) x London dummy	-0.747 (3.119)
LPA FEs	Yes
Year FEs	Yes
Observations	7,555
Number of LPAs	344
Kleibergen-Paap F	9.747

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. [a)] First stage results reported in columns (3) and (4) of Table B1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). [b)] PRP vs. market rent outliers (mean log market rent > 7.5, based on Figure 4) removed.

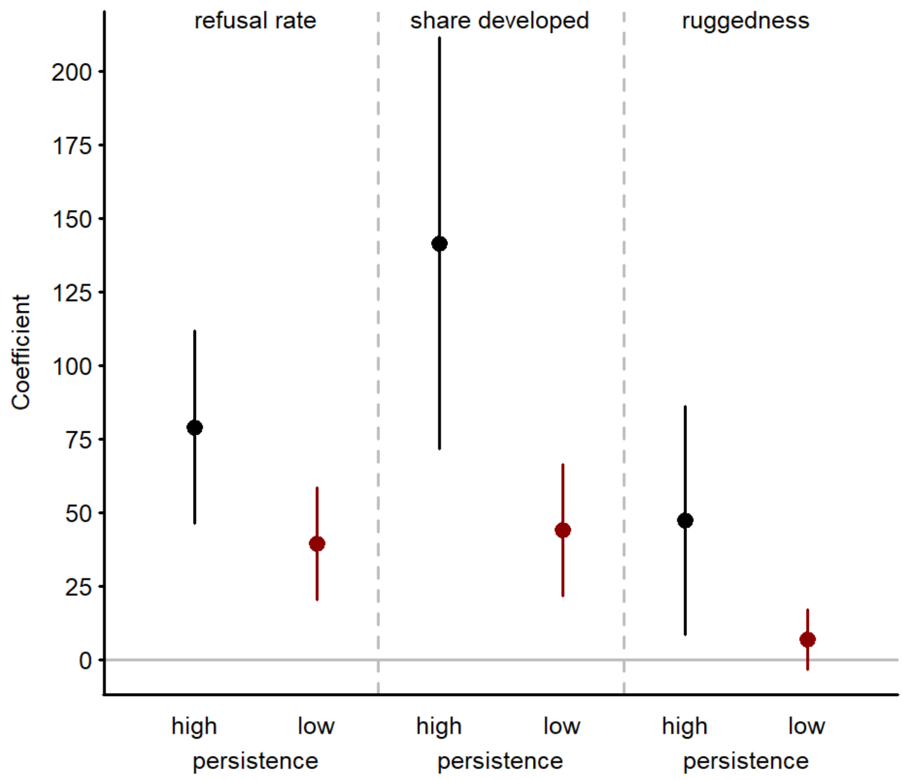
Heterogeneity in Persistence Across LPAs

According to our *Proposition 3 (iii)*, higher persistence of housing demand shocks can be expected to amplify the interaction effect of the demand shocks with the housing supply price elasticity on the price-to-rent ratio. Certain industries may be subject to shocks exhibiting greater persistence than others, resulting in differences in persistence across LPAs. To test the proposition, we split the sample at the median (see Figure 5) and estimate the baseline regression separately for each half of the sample. The estimated interaction effects are shown in Figure 6, with the three interaction terms separated by vertical dashed lines. The results lend support to *Proposition 3 (iii)*, with significantly stronger effects in high-persistence LPAs.

Positive vs. Negative Labor Demand Shocks

Recall from Section 2.3 that, because of the kinked nature of the supply curve, the theoretical predictions differ markedly, depending on whether local housing demand expands or contracts (*Proposition 4*). The results presented in Table 4 do not account for this distinction. To test *Proposition 4*, we therefore split the sample into LPA-years with positive and negative local housing demand shocks, as indicated by the year-to-year difference in the local labor demand measure. In the baseline sample from 1997 to 2018, there are 6,254 location-year observations with a positive and 1,248 with a negative labor demand shock. There are no locations that experienced negative labor demand shocks after 2015, which is why the Help to Buy dummy is not identified in column (2) of Table 5. Moreover, we restrict the two sub-samples to the same set of LPAs (excluding LPAs where local labor demand increased in every single year).

Figure 6
Interaction Effects for LPAs with Above- and Below-Median Persistence



Notes: The graph displays regression coefficients obtained from estimating equation (7) for LPAs with above- and below-median persistence in log labor demand shocks of 0.558, respectively, using only years with positive labor demand shocks (see the discussion relating to Table 5 below). The vertical bars indicate 95% confidence intervals (clustered by LPA).

We report the results in columns (1) and (2) of Table 5 (second stage) and Table B2 (first stage). Column (1) of Table 5 reveals that periods of positive local labor demand shocks are the main drivers behind the baseline results. All local labor demand-interaction terms, as well as the independent effect of this measure, are highly significant with the expected sign and (slightly) stronger than in the full sample. In contrast, when considering periods with declining local labor demand in column (2), the independent effect remains significant and gets larger in magnitude, while all three interaction terms are much closer to zero and no longer statistically significant, consistent with *Proposition 4*. Table B2 reveals that the excluded instruments again correlate strongly and in expected ways with the endogenous supply constraints. We present corresponding results for house prices and rents in Table O-B1 of Online Appendix O-B. The results are qualitatively similar.

Table 5

Separate Results for Periods with Positive and Negative Labor Demand Shocks

	(1)	(2)
	Price-to-rent ratio 2SLS [a)] 1997-2018 [b) c)] △LLD>0	Price-to-rent ratio 2SLS [a)] 1997-2018 [b)] △LLD≤0
Log(local labor demand)	34.730***	46.495***
	(13.266)	(13.665)
Av. refusal rate × log(local labor demand)	64.089***	23.993
	(9.927)	(15.169)
Share developed × log(local labor demand)	84.998***	7.529
	(19.589)	(10.140)
Altitude range × log(local labor demand)	26.218***	-1.405
	(9.889)	(3.704)
Help to Buy (post-2015) x London dummy [d)]	-1.587	
	(3.359)	
LPA FEs	Yes	Yes
Year FEs	Yes	Yes
Observations	6,254	1,248
Number of LPAs	341	341
Kleibergen-Paap F	9.001	6.985

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. [a)] First stage results are reported in Table B2. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). [b)] PRP vs. market rent outliers (mean log market rent > 7.5, based on Figure 4) removed. [c)] LPAs w/o negative local labor demand shocks removed to make the geographic extent of the sample (i.e., 341 LPAs) comparable. [d)] The Help to Buy dummy is not identified in column (2) because all locations experienced increasing local labor demand after 2015.

Rent Growth Expectations

An important assumption in our theoretical framework is that households form rational expectations about future rent growth (Assumption 3). Local labor demand shocks affect future rent growth expectations via autocorrelation in the shocks. Local supply constraints (at least for positive shocks) and the degree of local autocorrelation are important in that they amplify the impact of the initial shock on rent growth expectations. Conditional on the discount rate, house prices can then be expected to rise more strongly than current rents in response to a positive demand shock. This is because of the increase in future rent growth expectations.[26]

The results presented so far are consistent with this argument. However, in our main empirical analysis, we only observe shifts in local labor demand, a varying degree of local autocorrelation in that demand, and local supply constraints (i.e., the extent to which supply responds to rising demand). We do not observe rent growth expectations at the LPA-level. To test our proposed mechanism more directly, we therefore make use of a unique data set provided by the Royal Institute of Chartered Surveyors (RICS) that asks surveyors about their shorter-run (12 months) and longer-run (5 years) rent growth expectations, for the nine Government Office Regions of England. House purchase- and rent-decisions are ultimately made by households. However,

[26] Changing discount (or interest) rates are the other major factor affecting the price-to-rent ratio. One concern is that the impact of the changing discount rate on the price-to-rent ratio may vary locally. We address this concern in Section 3.5 below.

households in England rely heavily on the assessments provided by surveyors. The data are available from 2013 to 2018. We use the RPIX to deflate the rent growth expectations and geographically match this data at the regional level with our main panel. Panel E of Table 1 displays summary statistics for the one- and five-year expected real rent growth variables.

In Table 6 we show that shocks to local labor demand induce agents in the market to update both their one- and five-year rent growth expectations. Rent growth expectations adjust more strongly in locations characterized by tight local supply constraints and a higher persistence in demand shifts, consistent with our proposed theoretical mechanism. Moreover, the impact of persistence is greater on five-year rent growth expectations. The results are robust to adding region- and year-fixed-effects. We caveat that the number of observations is relatively small.

Table 6

Labor Demand Shocks and Rent Growth Expectations at Regional Level

	(1)	(2)	(3)	(4)	(5)	(6)
	One-year-ahead rent growth expectation			Five-year-ahead rent growth expectation		
	OLS	OLS	OLS	OLS	OLS	OLS
Change in log labor demand, ΔLLD	0.949***	0.947***	0.507	0.992***	0.992***	0.378
	(0.071)	(0.068)	(0.453)	(0.060)	(0.058)	(0.835)
Av. refusal rate × ΔLLD	0.212***	0.222***	0.214**	0.203**	0.211**	0.192**
	(0.059)	(0.059)	(0.068)	(0.084)	(0.083)	(0.080)
Share developed × ΔLLD	0.347***	0.351***	0.319***	0.254***	0.257***	0.182***
	(0.025)	(0.023)	(0.050)	(0.025)	(0.026)	(0.052)
Altitude range × ΔLLD	0.134**	0.141**	0.136**	0.104	0.108*	0.094
	(0.043)	(0.042)	(0.042)	(0.057)	(0.056)	(0.053)
High-persistence × ΔLLD	0.273*	0.281**	0.342**	0.498***	0.500***	0.547***
	(0.122)	(0.120)	(0.146)	(0.094)	(0.089)	(0.116)
Region FE	No	Yes	Yes	No	Yes	Yes
Year FE	No	No	Yes	No	No	Yes
Observations	54	54	54	54	54	54
Number of GORs	9	9	9	9	9	9
Adj. R-squared	0.388	0.404	0.905	0.379	0.433	0.888

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The outcome variable is the regional rent growth expectation, one and five years ahead. We use the change in the labor demand shock preceding the survey year. Expectations are based on a survey conducted by RICS, aggregated to Government Office Regions in England. We aggregate all remaining variables to the regional level. High-persistence regions have above-median persistence in demand shocks, where persistence is averaged over all LPAs in a region. The supply constraints are standardized to mean zero and standard deviation one.

3.5 Alternative Mechanisms

While our findings are consistent with our proposed mechanism, several alternative explanations are also conceivable. We explore these one by one below and report additional results as tables and figures in Appendices B and C.

Segmented Markets and Local Trends in Income Inequality

To the extent that owner-occupier and rental markets are segmented, and local income inequality is rising over time in a cyclical fashion, this too could explain a rising and cyclical price-to-rent ratio. Moreover, if this rise in income inequality were more pronounced in London than elsewhere, it could explain why the rise in the price-to-rent ratio has been most pronounced in the capital.

To explore this potential alternative mechanism and control for it, we draw on detailed income data at LPA-level that is available from 1997 onwards. We calculate the income dispersion as the log difference between the 80% and the 20% quantile of the local income distribution (male full-time earnings at workplace). Figure C1 displays the averages for England, London, the South East, and the North East over our sample period. There are no signs of divergence between London and the South East vis-à-vis England as a whole or the North East. If anything, income inequality *increased* slightly in the North East, but remained constant in the South East and London, suggesting that differential trends in income inequality may not explain the divergence of price-to-rent ratios between regions in England.

To test this conjecture more rigorously, we add this measure of local income inequality as a control to the baseline regression in Appendix Table B3, column (1). This hardly affects the coefficients of the log labor demand and its interactions with the supply constraints measures. When adding interactions of income inequality with the supply constraints in column (2), our main results are virtually unchanged. Moreover, the income inequality coefficients in the two specifications are mostly insignificant. As an alternative measure for income inequality, we employ an approximated Gini coefficient in columns (3) and (4), leading to very similar results.[27]

Finally, we replace the price variable in the calculation of the price-to-rent ratio by the average price for dwelling units (rather than using both single-family units and dwellings). Dwellings are more likely to be renter-occupied. Moreover, the bulk of rental properties in England is owned by private landlords, with the properties in question often being the owner's previous home. Hence, we would expect rental and owner-occupied dwellings to be much closer substitutes. The results in column (5) are again robust to this change, suggesting that prices in the different market segments co-move closely.

Financing Conditions and Idiosyncratic Risk

Unobserved shocks to the relative (financing-)cost of residential real estate could be correlated with changes in our measure of local labor demand, for instance due to financial innovation. Moreover, lower costs or higher availability of mortgage credit could induce higher demand for owner-occupied housing relative to renting. To the extent that housing supply is relatively price inelastic, we may then expect prices to increase relative to rents.

A fall in the real rate of mortgage interest or in the mortgage interest rate spread (i.e., the difference between the mortgage interest rate and the sight deposit rate) may make homeownership more desirable relative to (i) renting and (ii) other investment options.[28] This is a concern in our empirical setting to the extent that changes in the interest rate or the spread are correlated with changes in our labor demand measure. To address this, in column (1) of

[27] The Gini coefficient is based on the first to the eighth decile, the first and third quartile, and the mean of the local income distribution. We do not use the ninth decile because it has many missing values.

[28] We use the Bank of England's quoted mortgage interest rate deflated by the RPIX. Over our sample period, the real mortgage rate ranges between -1.22 and 6.02, while the spread ranges between 3.13 and 4.85, with standard deviations of 1.96 and 0.49, respectively.

Appendix Table B4 we add the real rate of mortgage interest interacted with the supply constraints (instrumented) as additional controls. In column (2), we repeat this exercise but use the spread interacted with the supply constraints (instrumented as well) instead.

Our main results are only marginally affected when we add these controls. We caveat that identification is weaker in these two regressions, as indicated by a comparably low Kleibergen-Paap F-statistic. Nonetheless, the estimates indicate that the real rate of mortgage interest and the mortgage interest rate spread interactions are quantitatively very substantially less important than the local labor demand interactions, suggesting that changes to the cost of mortgage financing cannot explain much of the large spatial variation in the price-to-rent ratio observed during our sample period. For instance, when we compare two locations that differ in their regulatory restrictiveness by one standard deviation, lowering the mortgage interest rate by one standard deviation (1.96) increases the difference in the price-to-rent ratio by only $1.96 \times 0.246 = 0.48$. In contrast, increasing the log labor demand by one within-standard deviation (0.05) has a much larger effect of $0.05 \times 56.1 = 2.81$. In a similar vein, decreasing the spread by one standard deviation (0.49 percentage points) increases the difference in the price-to-rent ratio by only $0.49 \times 0.59 = 0.29$, compared to 3.01 for a one-within-standard deviation increase of the log labor demand.

Since structures depreciate, while land does not, the investment horizons between these two components of housing are likely to differ. When the land value share is large, like in superstar cities, long-term real interest rates may be relatively more relevant for house price determination than the mortgage interest rate, which captures the financing of both land and structure. Moreover, land values make up a larger share of the overall property value in more developed places. To the extent that changes in the long-term real interest rate is also correlated with shocks to labor demand, this represents a threat to identification. In column (3), we address this concern by interacting the (standardized) land value share in 2000 by LPA with the long-term real interest rate.[29] The estimated coefficient is positive and significant. Our main findings are robust to adding this control.[30]

Finally, the reader might be concerned that the labor demand shock measure is correlated with local risk premia. Risk premia may change over time, e.g., due to changes in market liquidity. To address this concern, we construct a measure of local idiosyncratic price risk at the LPA-year level, using repeated sales[31] from the Land Registry (1995-2018) and closely following the methodology outlined in Giacoletti (2021). In addition, we construct an alternative measure based on repeated sales and the residual variation after controlling for housing unit and year fixed effects. We describe both methodologies in Online Appendix O-D. The correlations

[29] We use the long-term rate from the Bank of England's 'A Millennium of Macroeconomic Data' compendium, updated to 2018. To construct the land value share, we make use of data on land value per hectare in 2000 – the first year with available data – published by the Valuation Office Agency. We assume an average plot size of 100m² and divide the resulting land price by the price of an average home in 2000.

[30] When we instrument for the land value share using the historic population density in 1911, the results are unchanged, except that the coefficient of the long-term interest rate interacted with the land value share becomes negative and insignificant. When we repeat this exercise but use the long-term real interest rate interacted with the supply constraints (instrumented) instead, our main findings are again unaltered. The latter interaction captures any variable that could be correlated with local supply constraints, not just land value shares.

[31] Our baseline results are robust to using a repeated sales-based house price index instead of the constant-composition house price index that covers the entire sample period (1974-2018). Results are available on request.

between changes in the local labor demand measure and the idiosyncratic risk measures are very low, with 0.030 for the Giacoletti (2021) measure and 0.072 for the FE-based measure, strongly suggesting that within-LPA variation in idiosyncratic risk cannot explain our baseline results (which are identified by within-LPA variation in labor demand). When adding the idiosyncratic risk measures as a control to the baseline regression in columns (3) and (4) of Appendix Table B4, the coefficients of our main interaction effects remain very stable.[32]

Rent Stickiness in Existing Contracts

A fourth concern relates to the use of surveyed rents, which are derived from movers and stayers. These could be stickier than rents measured through online offers of vacant rental units, or from mover households alone. In institutional settings characterized by tenancy rent control, such measures can severely underestimate rent increases during housing booms. Comparable rules however do not exist in the English rental housing market, so that a landlord – in principle – can offer a new rental contract to her tenant each year. It could still be that landlords refrain from adjusting rents upwards, even in situations where local housing demand increases.[33] However, such behavior should become much less important over a longer time horizon, when more tenants have moved, and when the gap to the 'market rent' has widened, making a rent adjustment significantly more likely. We therefore consider regressions in one-, three-, and five-year differences as an alternative to the fixed effects approach. To account for differences in local average growth rates and average yearly changes, we also control for LPA- and year-fixed effects. The first column of Appendix Table B5 reveals that the results for one-year differences are very similar to the baseline results. When using three-year differences in column (2), the independent effect of the local labor demand shock becomes weaker and turns insignificant. The interaction effect of the local labor demand shock and the share developed land also gets somewhat weaker, but remains highly significant, while the interaction effect with the refusal rate gets larger. This pattern does not change much when using five-year differences in column (3). Overall, these results support the view that due to the institutional setting, rent stickiness in existing contracts is not an important phenomenon in England.

Global Investor Demand for Properties in London

Finally, we examine the hypothesis that global investor demand for London properties and other London-specific shocks may explain the relative increase of the price-to-rent ratio in London over our sample period. The regression residuals for Greater London put an upper bound to the quantitative importance of these channels. They should capture the overall impact of all other relevant factors orthogonal to the local labor demand shocks.

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[32] In column (3), the coefficient of the Giacoletti (2021) risk measure is positive and significant, while the FE-based measure in column (4) is negative and insignificant. The flipped sign mirrors the timing problem discussed in Online Appendix O-D. Both measures are negatively correlated with the price-to-rent ratio across LPAs, consistent with results from the finance literature, e.g., Amaral *et al.* (2021).

[33] In this setting, the relative bargaining power depends on the landlord's costs to fill a vacancy and on the tenant's moving costs (including the costs of renting another housing unit). In markets with increasing housing demand, it seems likely that vacancy risk is relatively low, whereas moving and search costs for the tenant may be substantial due to competition from other renters. This suggests that rent adjustments during a tenancy should be common during house price booms.

Panel A of Figure C2 clearly shows that there is little room for global ('out-of-town') investor demand as an explanation for the substantial increase of London's price-to-rent ratio. Between 1997 and 2003, the average residual in London was positive but small. It was negative from 2004 to 2008 and it has been hovering around zero since 2009. Overall, the net impact of other London-specific factors seems to be rather small.[34]

Panels B to D display analogous graphs for the South East (another region with a white-collar service-oriented workforce), the North East, and England as a whole. The predicted and actual price-to-rent ratios are reasonably close in all cases, suggesting that region-specific global investor demand or other region-specific factors may not have been driving forces explaining the regional divergence in the price-to-rent ratios since 1997.

3.6 Robustness Checks

In this section, we explore several empirical concerns and test the robustness of our baseline results along these dimensions. We report results as tables and figures in Appendices B and C.

Selection of Instrumental Variables

A first concern is that our estimated coefficients of interest may be sensitive to the choice of instrumental variables used to identify the refusal rate of major residential planning applications. In our baseline specification, we employ three separate instrumental variables jointly: the share of greenbelt land in 1973, the change in the delay rate, and the vote share of the Labour party in the 1983 General Election. Appendix Table B6 reports results for six different alterations of the baseline specification (Table 4). The first three models drop one instrument at a time. Specifications (4) to (6) then report estimates keeping only one of the three instruments at a time. The coefficients of interest remain stable across all six specifications, with the Kleibergen-Paap F-statistic varying more markedly but generally indicating that weakness of identification is not a concern.

Choice of Rent Measure and Sample Restrictions

A second concern is that the PRP rental data used to calculate the price-to-rent ratio may not adequately reflect the behavior of market rents. We use PRP rents in the first place because it enables us to extend the study period to 22 years, covering nearly two full local housing market cycles. While the correlation between log PRP rents and log market rents is very strong (0.86), as Figure 4 illustrates, our full sample of LPAs contains several (high-end market) outliers with a somewhat weak relationship between PRP rents and market rents. Here, we test whether our results are robust to (i) using a different approach to selecting LPAs and (ii) using the full sample of market rents instead of PRP rents. At a basic level, PRP rents are a good proxy for market rents in our empirical setting if their year-to-year correlation within an LPA is sufficiently strong. Appendix Figure C3 depicts a kernel density plot of the correlation between the change in PRP rents and the change in market rents at LPA-level. In most LPAs, the

[34] This does not preclude that global investor demand is an important driver of local house prices in specific market segments, such as the prime market in central London (Badarinza and Ramadorai 2018) or in Manhattan in New York City (Favilukis and Van Nieuwerburgh 2021). However, in the case of London, these prime markets or neighborhoods are too small to markedly influence the price development in the entire Greater London area.

correlation is positive, or even strongly positive. However, there are also some LPAs where the correlation is weak or even negative.

In Appendix Table B7, we restrict the sample based on Appendix Figure C3. A natural threshold is at zero, and we test two further thresholds based on the two local minima of the density graph at 0.1 and 0.45, respectively. In each case, we restrict the sample to LPAs that lie to the right of the threshold, see columns (1) to (3).[35] The interaction coefficients are somewhat larger than in the baseline specification, and the independent effect of the local labor demand measure is smaller and insignificant.

Column (4) reveals that our main results are also robust towards using market rents for the calculation of the price-to-rent ratio and to using the full sample of 353 LPAs. Since market rents are only available from 2010 onwards, we re-estimate the baseline regression based on PRP rents in column (5), for the sub-sample starting in 2010, leading to the same pattern of coefficients as in column (4). Overall, these results strongly suggest that PRP rent dynamics are very similar to the dynamics of market rents, at least along the dimensions we consider in this analysis.

Price-to-Rent Ratio in Logs

Third, one might be concerned about the functional form of the outcome variable. Price-to-rent ratios in levels are not necessarily normally distributed, with potentially large outliers that may exert strong influence on regression coefficients. Taking the log price-to-rent ratio reduces greatly the impact of outliers.[36] In column (1) of Appendix Table B8, we replicate the baseline specification from Table 4, column (1). In column (2), we do the same, but use market rents, replicating the specification reported in Table B7, column (4). In both cases, the labor demand interactions with the supply constraints are quantitatively similarly important as in our baseline specification.

Local Labor Demand Shock: A Placebo Test

A fourth concern is that the initial industry composition used for the construction of the shift-share measure could correlate with unobserved shocks to the relative attractiveness of renting versus owning. This concern relates to the interpretation of the shift-share instrument as a weighted sum of generalized difference-in-differences estimators, where each estimator builds on a comparison of initial employment shares in a particular industry (Goldsmith-Pinkham *et al.* 2020). In this interpretation, endogeneity concerns arise from correlations between changes in unobserved confounders and the initial industry composition. While our setting differs from that discussed in Goldsmith-Pinkham *et al.* (2020) – most importantly because in our setting the impact of labor demand shocks is heterogeneous across space and over time, but also because the industry shares pre-date our sample period by more than 25 years – we can explore the degree to which our results depend on the initial industry composition *alone*.

With endogenous initial industry shares, the regression coefficients could be positive and significant even when creating the shift-share instrument from any other set of serially

[35] Our results are similar (with coefficients being somewhat larger) when we use the full sample of 353 LPAs.

[36] Log price-to-rent ratios have a less straightforward interpretation. We therefore use the price-to-rent ratio in levels for our main analysis.

correlated time series. To test this, we re-create the shift-share measure based on simulated employment series for the seven industries. We assume that the national-level time series are autocorrelated processes of order p and we select p by the Akaike information criterion.[37] We then simulate the seven industry time series and create the shift-share measure based on the actual industry composition and the simulated time series to get a placebo-measure of local labor demand. With this placebo measure, we then estimate the baseline model. We repeat the whole exercise 2000 times to get a parameter distribution for each regression parameter of the baseline model. If the initial industry composition were exogenous to the model, we would expect that these distributions center on zero, and that our baseline estimates are located towards the right tails of the distributions. Appendix Figure C4 displays the coefficient distributions for the independent effect of the local labor demand measure and its three interaction terms with supply constraints. All estimated baseline coefficients are near or beyond the right tail of the respective simulated coefficient distribution.

Adjusted Local Labor Demand Shock Measure (Excluding Banking and Real Estate)

A fifth and related concern is that local labor demand shocks could also affect local credit availability. This would obfuscate the impact of shocks to overall housing demand on the price-to-rent ratio, due to the direct and distinct impact of credit supply on the relative attractiveness of owning versus renting. In Appendix Table B9, we therefore replace the original labor demand-measure with an adjusted version: The labor demand measure relies on time-series variation of employment in seven industries, one of them being the services and distribution sector. Two sub-sectors are banking and real estate services. We replace the employment series for the services and distribution sector by an adjusted series that excludes the two sub-sectors. We then recreate the shift-share labor demand measure using this adjusted series. Our results of interest hardly change, suggesting that shocks to employment in the banking and real estate services sectors do not influence our findings.

4 Quantitative Analysis

To assess the quantitative importance of the mechanism we uncover, we decompose the predicted evolution over time of the price-to-rent ratio into its aggregate (macro) component and its local components (impact of local labor demand shocks interacted with the housing supply constraints). Second, we conduct a counterfactual analysis where we compare the predicted price-to-rent ratio in selected regions, to the price-to-rent ratios of two hypothetical locations with average and relatively lax housing supply constraints, respectively.

4.1 Decomposition

In Panel A of Figure 7, we use the coefficients from columns (1) and (2) of Table 5 – depending on whether an LPA was hit by a positive or negative labor demand shock in a given year – to decompose the overall evolution of the predicted price-to-rent ratio (blue dashed line) in Greater London (consisting of 32 LPAs), into the impact of the aggregate component (the year fixed effects including the impact of Help-to-Buy; red dashed line), and into the effects of local labor demand shocks in conjunction with the housing supply constraints (the difference between the

[37] The Akaike information criterion selects a lag order of 2 for the construction industry, and a lag order of 1 for all other industries.

red and the blue dashed lines). In addition, it shows the actual price-to-rent ratio (solid black line). We select London because it experienced strong labor market shocks and has severely constrained housing supply, mainly due to a high share of developed land. In Online Appendix Figure O-C1, we show corresponding results for the neighboring South East, a region that is characterized by very tight regulatory constraints, and for the North East, a region with comparably lax supply constraints. Both London and the South East are good examples of "location B", while the North East is a good example of "location A" in Figure 3.

Panel B displays the corresponding log labor demand. In years with dark shading, all LPAs of Greater London experienced decreasing demand. Light-grey shading indicates periods where at least one LPA experienced a negative demand shock. In periods without shading, demand increased in all LPAs. Panel C displays the Repo/Official Bank Rate[38] (light grey line), the real mortgage interest rate (dashed red line), and the real long-term interest rate (short-dashed blue line). Falling real interest rates are a potential contributor to the year fixed effects, and a plausible explanation for rising aggregate price-to-rent ratios during our sample period.

In Panel A, the 'unexplained component' captured by the year fixed effects increased strongly between 1997 and 2004, but remained stable from 2004 to 2010, and then fell substantially to later recover slightly between 2010 to 2018. Panel C suggests that interest rates may have been a substantial factor in explaining the increase in the price-to-rent ratio between 1997 and 2004 – a period where all three reported interest rates decreased substantially. The time fixed effects explain 62.8% of the overall increase up to the start of the Great Financial Crisis in 2007. However, real interest rates may not explain the sharp decline in the price-to-rent ratio that ensued during the Great Financial Crisis years, a period where all three reported interest rates decreased significantly. The time fixed effects also cannot explain the sharp rise in the price-to-rent ratio since 2010.

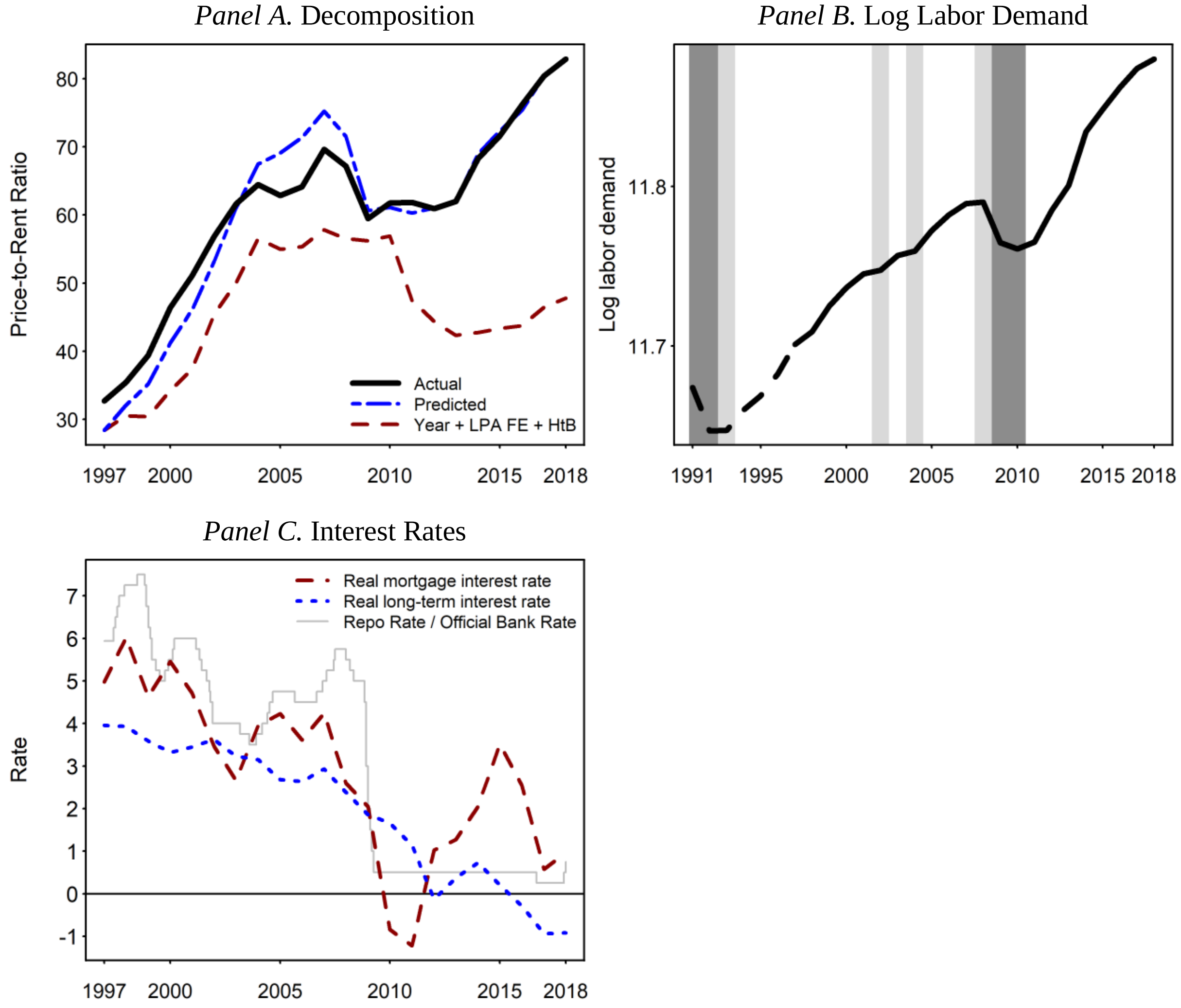
The interaction of labor demand and housing supply constraints in the tightly supply-constrained Greater London contributed significantly to the marked increase in the price-to-rent ratio up to 2007, leading to a considerable gap between the red and blue lines. The total effect of local labor demand shocks and their interactions with supply constraints represent 37.2% of the overall increase between 1997 and 2007. More importantly, the labor demand shock × supply constraint-interactions can (more than) fully explain the decline in the price-to-rent ratio between 2007 and 2010, as well as the ensuing increase in the ratio between 2010 and 2018.

Overall, the decomposition strongly suggests that differential rent growth expectations, driven by persistence in the labor demand shocks and inelastic housing supply, are quantitatively important in explaining the price-to-rent ratio also over extended periods of time. In fact, when comparing 1997 to 2018, 64.3% of the increase in the price-to-rent ratio in London can be explained by persistent labor demand shocks in conjunction with local supply constraints.

[38] The Official Bank Rate was introduced in 2006 and replaced the Repo Rate in use since 1997.

Figure 7

Decomposition of the Price-to-Rent Ratio and Potential Drivers in London



Notes: Panel A displays the actual (solid black line) and predicted (dashed blue line) price-to-rent ratio, as well as the evolution of the price-to-rent ratio that is attributed to the fixed effects and Help-to-Buy (dashed red line), based on the two models in Table 5. The models were used to compute LPA-level predictions, that were aggregated to the Government Office Region of London, employing the number of households in each LPA in 2011 (Census) as weights. Panel B displays the corresponding labor demand variable, aggregated to London. In years with dark shading, all LPAs experienced decreasing demand. Light-grey shading indicates periods where at least one LPA experienced decreasing demand. In periods without shading, demand increased in all LPAs. Panel C displays the Repo/Official Bank Rate (light grey line), the real mortgage interest rate (dashed red line), and the real long-term interest rate (short-dashed blue line); data source: Bank of England, A Millennium of Macroeconomic Data.

Figure O-C1 displays the respective decompositions and labor demand series for the South East and the North East of England. The decomposition for the South East is qualitatively similar to that for London. However, the picture is reversed for the North East. In fact, during the most recent boom period, the aggregate ‘unexplained’ component grew more strongly than the actual and predicted price-to-rent ratios, implying that the labor demand shocks had an attenuating effect on the price-to-rent ratio, consistent with the case of ‘Location A’ in Figure 3.

4.2 Decomposition of the Labor Demand and Supply Constraints Interactions

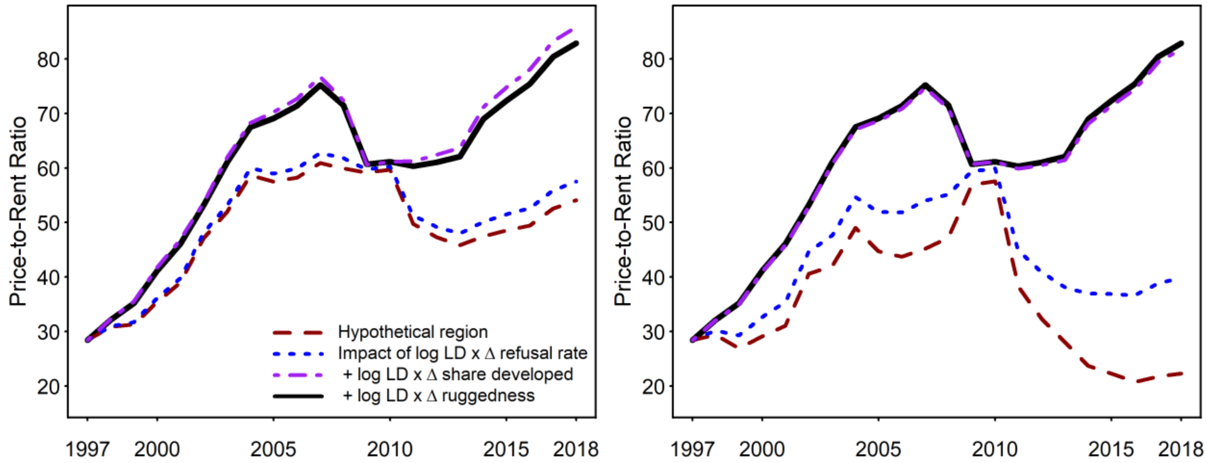
In Figure 8, we decompose the impact of the labor demand shock in London further into the contributions of the different interaction terms with the housing supply constraints. In Panel A of Figure 8, we compare London to an average LPA in England, assuming both London and the hypothetical average location were hit by the same labor demand shocks. The graph reveals that the cyclical differences between London and the hypothetical location with average supply constraints are fully explained by the interaction of labor demand shocks with housing supply constraints, the main effect coming from the interaction with the share of developed land. The altitude range and regulatory restrictiveness play a much smaller role in this case, simply because the averages of these variables are similar in London and England as a whole (with the refusal rate only being slightly above average, and ruggedness slightly below).

Figure 8

Decomposing the Difference in Price-to-Rent Ratios between London and Other Locations

Panel A. Comparison to a Location with Average Supply Constraints

Panel B. Comparison to a Location with Lax Supply Constraints



Notes: Both graphs are based on the models displayed in Table 5. The models were used to compute LPA-level predictions, that were aggregated to Government Office Regions, employing the number of households in each LPA in 2011 (Census) as weights. Panel A compares the prediction for London (black solid line) to the prediction for a hypothetical location with average supply constraints (dark-red dashed line) and decomposes the difference into the impact of the local labor demand interactions with the differences in local regulatory restrictiveness (blue dotted line), the share developed land (purple dashed-dotted line), and ruggedness (the difference between the purple dashed-dotted line and the black solid line). Panel B repeats this exercise for a hypothetical location with all three supply constraints at the respective 10% quantile.

Arguably, the English planning system is one of the strictest – perhaps the strictest – in the world. Consequently, the average location in our sample is likely a tightly regulated place by international standards. Moreover, in comparison to the United States and other countries with vast amounts of open land, England's population density is high. Both factors suggest that the decomposition exercise in Panel A of Figure 8 understates the importance of local housing supply constraints relative to countries with a higher average housing supply elasticity.

We therefore conduct an additional decomposition exercise in Panel B of Figure, using a hypothetical comparison region that exhibits rather lax supply constraints. We define this region by taking the first decile of each supply constraint-variable (refusal rate, share developed, and elevation range). The decomposition exercise shows the impact of changing the supply constraints from the first decile to London's level of supply constraints. The effect works through the interaction with the demand shocks, which we again assume to be the same for both locations.

The cyclical differences between the hypothetical region with lax supply constraints and London are much wider. Remarkably, they disappear entirely during the Great Financial Crisis – consistent with theory – and reappear in the ensuing boom. Relative to the first decile-level of supply constraints, London exhibits a substantial degree of regulatory restrictiveness, which leads to a much larger impact of the refusal rate. The share developed measure is again the most important variable for explaining the large cyclical swings.

The graph also suggests that the price-to-rent ratio would have decreased slightly over our sample period if housing supply constraints in England were as lax as in the hypothetical location. Our empirical and theoretical models attribute this decrease to the impact of persistent shocks to aggregate housing demand in conjunction with lax supply constraints in the hypothetical region, which may counteract the effects of declining real interest rates and improving financing conditions.

We repeat the exercise in Online Appendix Figure O-C2 for the South East in Panels A and B, and the North East in Panels C and D. In the South East, regulatory restrictiveness is the most important driver. The North East has below-average regulatory restrictiveness and a below-average share of developed land, so that the interaction effects with these two variables have an attenuating effect on the price-to-rent ratio.

5 Conclusions

The underlying causes of the housing affordability crisis have been one of the most hotly contested debates in urban economics, but the topic has also raised interest among macroeconomists and financial economists. One question is particularly policy relevant: To what extent is the rising house price-to-rent ratio consistent with housing supply shortages?

In this study we provide a simple theory – tight supply constraints in conjunction with serially correlated demand shocks – to explain why (i) the increase in the price-to-rent ratio tends to be most pronounced in the most desirable and supply-constrained (superstar) cities of a country, (ii) the evolution of the price-to-rent ratio over time varies dramatically across locations within country, (iii) the price-to-rent ratio is cyclical in nature, and (iv) the price-to-rent ratio falls in markets (such as Japan) hit by prolonged negative demand growth.

Our empirical findings help to reconcile the mainstream urban economic and macroeconomic views: In line with the former view, our analysis highlights the importance of local long-run supply constraints – including regulatory constraints – in explaining why housing affordability has declined dramatically in superstar cities like London (and other thriving places) over the last two decades and why house prices in these places have risen even more strongly than rents. In line with the latter view, our analysis suggests that, at the aggregate level, when excluding a country's most thriving locations, *macroeconomic factors*, as summed up by the year fixed effects, haven been crucial drivers explaining the price-to-rent-ratio dynamics, especially during the run-up to the Great Financial Crisis. The year fixed effects are a 'black box' that are likely to capture changing financing conditions as well as aggregate supply constraints in conjunction with serially correlated aggregate housing demand shocks. Unpacking this black box is an intriguing and important question for future research.

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Appendices

Appendix A: Proof of Proposition 3

The first part of Proposition 3 is clear by inspection of the relevant expression in the main text. For parts (ii) and (iii), consider the price-to-rent ratio $Q = 1 + r\frac{E[R_2]}{R_1}$, and take the derivative w.r.t. the housing demand shock, ε:

$$Q_\varepsilon = \frac{\partial Q}{\partial \varepsilon} = r\frac{R_0(\beta\phi\delta + 1)(\beta\gamma\phi\delta + \beta\phi(\delta - 1) + \gamma)}{(\beta\phi + 1)(R_0(\beta\phi\delta + 1) + \varepsilon)^2}.$$

At $\varepsilon = 0$, this simplifies to

$$Q_\varepsilon|_{\varepsilon=0} = r\frac{\beta\gamma\phi\delta + \beta\phi(\delta - 1) + \gamma}{\beta\phi + 1}.$$

Taking the derivative w.r.t. β,

$$\frac{\partial}{\partial\beta}Q_\varepsilon|_{\varepsilon=0} = (\delta - 1)\frac{r\phi(1+\gamma)}{(1+\phi\beta)^2} < 0,$$

because $\delta < 1$ and all parameters are strictly positive. This shows part (ii) of Proposition 3.

Clearly, $\frac{\partial^2}{\delta\gamma\partial\beta}Q_\varepsilon|_{\varepsilon=0} < 0$, i.e., higher persistence amplifies the effect. This shows part (iii).

Appendix B: Appendix Tables

Table B1

First Stage Regressions relating to Table 2

	(1)	(2)	(3)	(4)
	Model (2) Refusal rate	Model (2) %Developed	Models (3), (4) Refusal rate	Models (3), (4) %Developed
Log(local labor demand, LLD)	0.077	0.138**	0.098	0.198**
	(0.060)	(0.062)	(0.075)	(0.080)
Altitude range × log(LLD)	-0.067	-0.392***	-0.066	-0.336***
	(0.052)	(0.041)	(0.046)	(0.034)
Change in delay rate × log(LLD)	-0.080*	0.017	-0.082*	-0.014
	(0.043)	(0.046)	(0.042)	(0.039)
Share Labour vote in 1983 × log(LLD)	-0.512***	0.245***	-0.588***	0.277***
	(0.070)	(0.050)	(0.041)	(0.041)
Share greenbelt in 1973 × log(LLD)	0.289***	0.008	0.270***	0.016
	(0.039)	(0.032)	(0.039)	(0.028)
Population density in 1911 × log(LLD)	-0.155*	0.432***	-0.010	0.537***
	(0.085)	(0.044)	(0.046)	(0.124)
Help to Buy (post-2015) x London dummy	0.058***	0.140***	0.032**	0.102***
	(0.017)	(0.017)	(0.013)	(0.012)
Observations	15,885	15,885	7,555	7,555
Number of LPAs	353	353	344	344
R-sq. overall	0.437	0.561	0.466	0.515
R-sq. within	0.434	0.655	0.465	0.555
R-sq. between	0.437	0.561	0.463	0.514

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Models (3) and (4) of Table 2 both have the same first stage. The excluded instruments are the change in delay rate, the share Labour vote in 1983, the share greenbelt in 1973, and population density in 1911 all interacted with the log(LLD).

Table B2

First Stage Regressions Relating to Models (1) and (2) of Table 5

	(1)	(2)	(3)	(4)
	Model (1) Refusal rate	Model (1) %Developed	Model (2) Refusal rate	Model (2) %Developed
Log(local labor demand)	0.128	0.266***	-0.255***	-0.070
	(0.089)	(0.094)	(0.092)	(0.105)
Altitude range × log(local labor demand)	-0.069	-0.360***	-0.044	-0.118***
	(0.049)	(0.037)	(0.056)	(0.018)
Change in delay rate × log(local labor demand)	-0.076*	-0.007	-0.082	0.014
	(0.043)	(0.041)	(0.082)	(0.027)
Share Labour vote in 1983 × log(local labor demand)	-0.603***	0.284***	-0.444***	0.169***
	(0.044)	(0.044)	(0.111)	(0.053)
Share greenbelt in 1973 × log(local labor demand)	0.270***	0.004	0.112	0.130**
	(0.039)	(0.028)	(0.071)	(0.056)
Population density in 1911 × log(local labor demand)	-0.008	0.525***	-0.084	0.728***
	(0.047)	(0.121)	(0.087)	(0.262)
Help to Buy (post-2015) × London dummy [a)]	0.031**	0.101***		
	(0.013)	(0.011)		
Observations	6,254	6,254	1,248	1,248
Number of LPAs	341	341	341	341
R-sq. overall	0.464	0.515	0.452	0.421
R-sq. within	0.472	0.551	0.358	0.827
R-sq. between	0.466	0.515	0.451	0.424

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The excluded instruments are the change in delay rate, the share Labour vote in 1983, the share greenbelt in 1973, and population density in 1911 all interacted with the log(local labor demand). [a)] The Help to Buy dummy is not identified in columns (3) and (4) because all locations experienced expanding local labor demand after 2015.

Appendix Table B3

Local Income Inequality and Market Segmentation

	(1)	(2)	(3)	(4)	(5)
	Price-PRP rent ratio, P80/P20	Price-PRP rent ratio, P80/P20	Price-PRP rent ratio, approx. Gini	Price-PRP rent ratio, approx. Gini	Dwelling Price-PRP rent ratio
Log(local labor demand)	39.520***	37.480***	36.158***	36.106***	29.713***
	(10.995)	(11.353)	(11.501)	(11.250)	(8.626)
Average refusal rate × log(local labor demand)	61.853***	62.339***	62.103***	61.362***	47.060***
	(9.268)	(9.189)	(9.580)	(9.034)	(6.983)
Share developable land × log(local labor demand)	80.630***	81.969***	81.789***	80.620***	63.988***
	(18.024)	(17.838)	(18.247)	(16.819)	(13.712)
Altitude range × log(local labor demand)	22.435***	22.731***	22.119**	21.711***	21.233***
	(8.486)	(8.478)	(8.607)	(8.116)	(6.703)
Help to Buy (post-2015) x London dummy	-1.101	-1.470	-1.263	-1.153	1.002
	(3.195)	(3.265)	(3.213)	(3.007)	(2.496)
Local income inequality	-1.744	-2.437**	1.490	2.766	
	(1.094)	(1.100)	(3.573)	(3.579)	
Average refusal rate × local income inequality		0.527		1.869	
		(1.354)		(5.322)	
Share developable land × local income inequality		-3.941**		18.229***	
		(1.757)		(6.829)	
Altitude range × local income inequality		1.091		-1.087	
		(1.404)		(5.198)	
LPA FEs	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes
Observations	6,830	6,830	6,735	6,735	7,555
Number of LPAs	344	344	344	344	344
Kleibergen-Paap F	11.07	5.30	10.71	6.31	9.75

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). In columns (1) and (2), local income inequality is the log difference between the 80% and the 20% quantile of the local earnings distribution. In columns (3) and (4), local income inequality is measured as the Gini coefficient (approximated from data on eleven quantiles across the local earnings distribution and the mean of the local distribution). Higher values of the local income inequality measure indicate greater inequality in both cases. The data source is the Annual Survey of Hours and Earnings, Table 7.1a - Weekly pay for full-time male workers at workplace.

Appendix Table B4

Mortgage Financing Conditions and Idiosyncratic Investment Risk

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	Price-PRP-rent ratio				
Additional controls:	Real Mortgage Rate × Supply Constraints	Mortgage Rate Spread × Supply Constraints	Long-Term Real Rate × Land Value Share	Idiosync. Risk (Giacoletti 2021)	Idiosync. Risk (OLS-FE residuals)
Log(local labor demand)	32.881***	30.032**	40.364***	38.501***	39.445***
	(12.219)	(12.514)	(11.093)	(10.736)	(10.676)
Av. refusal rate × log(local labor demand)	56.115***	61.115***	68.447***	60.503***	60.869***
	(7.658)	(8.376)	(12.620)	(8.968)	(9.035)
Share of developable land × log(local labor demand)	75.398***	85.019***	91.342***	81.029***	79.977***
	(15.843)	(16.569)	(23.764)	(18.202)	(18.170)
Altitude range × log(local labor demand)	18.760**	21.447***	23.228**	23.369***	23.033***
	(7.639)	(8.056)	(9.140)	(8.653)	(8.631)
Help to Buy (post-2015) x London dummy	-0.569	-2.080	-0.401	-1.864	-0.702
	(3.039)	(2.916)	(3.235)	(3.166)	(3.117)
Av. refusal rate × real mortgage rate	-0.246***				
	(0.089)				
Share of developable land × $r$eal mortgage rate	-0.186				
	(0.116)				
Altitude range × real mortgage rate	-0.220***				
	(0.061)				
Av. refusal rate × mortgage rate spread		-0.600			
		(0.401)			
Share of developable land × mortgage rate spread		0.911*			
		(0.474)			
Altitude range × mortgage rate spread		-0.839***			
		(0.289)			
Land value share × long-term real interest rate			0.845**		
			(0.423)		
Idiosyncratic volatility (Giacoletti 2021)				25.657***	
				(3.181)	
Idiosyncratic volatility (OLS-FE residual variation)					-9.060
					(6.378)
LPA FEs	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes
Observations	7,555	7,555	7,555	7,555	7,555
Number of LPAs	344	344	344	344	344
Kleibergen-Paap F	5.38	5.23	6.02	9.75	9.70

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). The interactions of the supply constraints with the real mortgage interest rate in column (1) and the spread (mortgage rate minus sight deposit rate) in column (2) are instrumented by the interactions of the respective variable with the instruments discussed in Section 3.2. The idiosyncratic volatility measure used in column (4) is based on Giacoletti (2021). The measure used in column (5) is based on residual variation at LPA-level, see Online Appendix O-D.

Appendix Table B5

Regressions in Differences

	(1)	(2)	(3)
	Δ Price-PRP rent ratio	Δ Price-PRP rent ratio	Δ Price-PRP rent ratio
	1-Year Diffs	3-Year Diffs	5-Year Diffs
Δ Log(local labor demand)	49.488***	11.579	0.470
	(9.016)	(18.291)	(29.523)
Av. refusal rate of major residential projects × Δ log(local labor demand)	65.163***	90.352***	87.443***
	(6.169)	(11.514)	(12.845)
Share of developable land developed in 1990 × Δ log(local labor demand)	53.198***	38.729***	37.168***
	(7.215)	(11.056)	(14.314)
Range between highest and lowest altitude × Δ log(local labor demand)	11.943***	11.477*	9.409
	(4.396)	(6.913)	(7.820)
Δ Help to Buy (post-2015) x London dummy	0.629	0.733	0.825
	(0.647)	(0.798)	(0.964)
LPA FEs	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes
Observations	7,211	6,523	5,835
Number of LPAs	344	344	344
Kleibergen-Paap F	13.33	14.51	13.81

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). All regressions are in differences. The column heading indicates the number of years over which the differences are computed (1, 3, and 5 years). The regressions also include LPA and year FEs to capture average LPA-level changes and national-level changes over the respective period.

Appendix Table B6

Robustness of Baseline Results to the Selection of Instrument

	(1)	(2)	(3)	(4)	(5)	(6)
	Excluding greenbelt instrument	Excluding delay rate instrument	Excluding Labour votes instrument	Only greenbelt instrument	Only delay rate instrument	Only Labour votes instrument
Log(local labor demand)	36.429***	38.278***	48.328***	46.009***	55.174***	33.821***
	(11.322)	(10.770)	(13.402)	(14.143)	(18.402)	(11.709)
Av. refusal rate of major residential	65.038***	62.086***	48.848***	52.034***	39.026*	69.331***
projects × log(local labor demand)	(13.285)	(8.927)	(10.704)	(11.122)	(22.928)	(14.161)
Share of developable land developed	84.418***	81.511***	79.825***	80.872***	74.630***	89.172***
in 1990 × log(local labor demand)	(22.250)	(18.045)	(17.349)	(17.170)	(21.323)	(23.060)
Range between highest and lowest	25.101**	23.763***	22.262***	22.878***	19.585*	27.256**
altitude × log(local labor demand)	(10.324)	(8.635)	(8.032)	(7.995)	(10.039)	(10.750)
Help to Buy (post-2015) × London	-1.777	-1.192	-0.699	-0.938	0.408	-2.727
dummy	(4.058)	(3.144)	(2.990)	(2.940)	(3.935)	(4.245)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Observations	7,555	7,555	7,555	7,555	7,555	7,555
Number of LPAs	344	344	344	344	344	344
Kleibergen-Paap F	7.00	13.36	17.35	23.01	5.78	10.06

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). The specifications use different sets of instruments for the average refusal rate, as denoted by the column headings.

Appendix Table B7

Robustness Checks for Selection of Rent Measure

	(1)	(2)	(3)	(4)	(5)
Dependent variable		Price- PRP rent ratio		Price- market rent ratio	Price- PRP rent ratio
LPA-level correlation of Δ PRP rent and Δ market rent	$>$ 0	$>$ 0.1	$>$ 0.45	-	-
Log(local labor demand)	28.310	31.014	9.438	-63.682**	-133.187**
	(19.739)	(19.960)	(37.545)	(24.840)	(57.563)
Av. refusal rate of major residential	90.957***	93.792***	99.873***	30.061***	72.496***
projects × log(local labor demand)	(18.412)	(19.699)	(35.269)	(4.347)	(8.133)
Share of developable land developed	156.620***	156.287***	180.748***	41.828***	95.139***
in 1990 × log(local labor demand)	(36.844)	(36.981)	(60.064)	(7.847)	(15.932)
Range between highest and lowest	50.043***	50.644***	68.564**	6.231*	6.697
altitude × log(local labor demand)	(16.448)	(16.915)	(28.261)	(3.322)	(7.162)
Help to Buy (post-2015) × London	-11.808**	-11.889**	-15.703*	-2.144***	-3.975*
dummy	(5.823)	(5.831)	(9.503)	(0.822)	(2.056)
LPA FEs	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes
Sample years	1997-2018	1997-2018	1997-2018	2010-2018	2010-2018
Observations	6,851	6,411	3,375	3,177	3,096
Number of LPAs	312	292	154	353	344
Kleibergen-Paap F	19.28	17.97	11.42	23.19	7.44

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). The specifications in columns (1) to (3) use different sub-samples, based on lower bounds for the correlation between changes in PRP rents and market rents at LPA-level. In column (4), the rent measure is based on market rents published by the Valuation Office Agency.

Appendix Table B8

Log Price-Rent Ratio as Outcome

	(1)	(2)
	Log Price-PRP rent, baseline sample	Log price-market rent all LPAs 2010-18
Log(local labor demand)	-0.089	-1.552**
	(0.184)	(0.640)
Av. refusal rate of major residential projects × log(local labor demand)	0.580***	0.826***
	(0.121)	(0.109)
Share of developable land developed in 1990 × log(local labor demand)	0.607***	1.118***
	(0.231)	(0.176)
Range between highest and lowest altitude × log(local labor demand)	0.079	0.068
	(0.114)	(0.080)
Help to Buy (post-2015) × London dummy	0.084**	-0.043**
	(0.042)	(0.021)
LPA FEs	Yes	Yes
Year FEs	Yes	Yes
Sample years	1997-2018	2010-2018
Observations	7,555	3,177
Number of LPAs	344	353
Kleibergen-Paap F	9.747	23.19

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The excluded instruments are the change in delay rate, the share Labour vote in 1983, the share greenbelt in 1973, and population density in 1911 all interacted with the log(local labor demand).

Appendix Table B9

Adjusted Labor Demand Shock (w/o Banking & Real Estate Services)

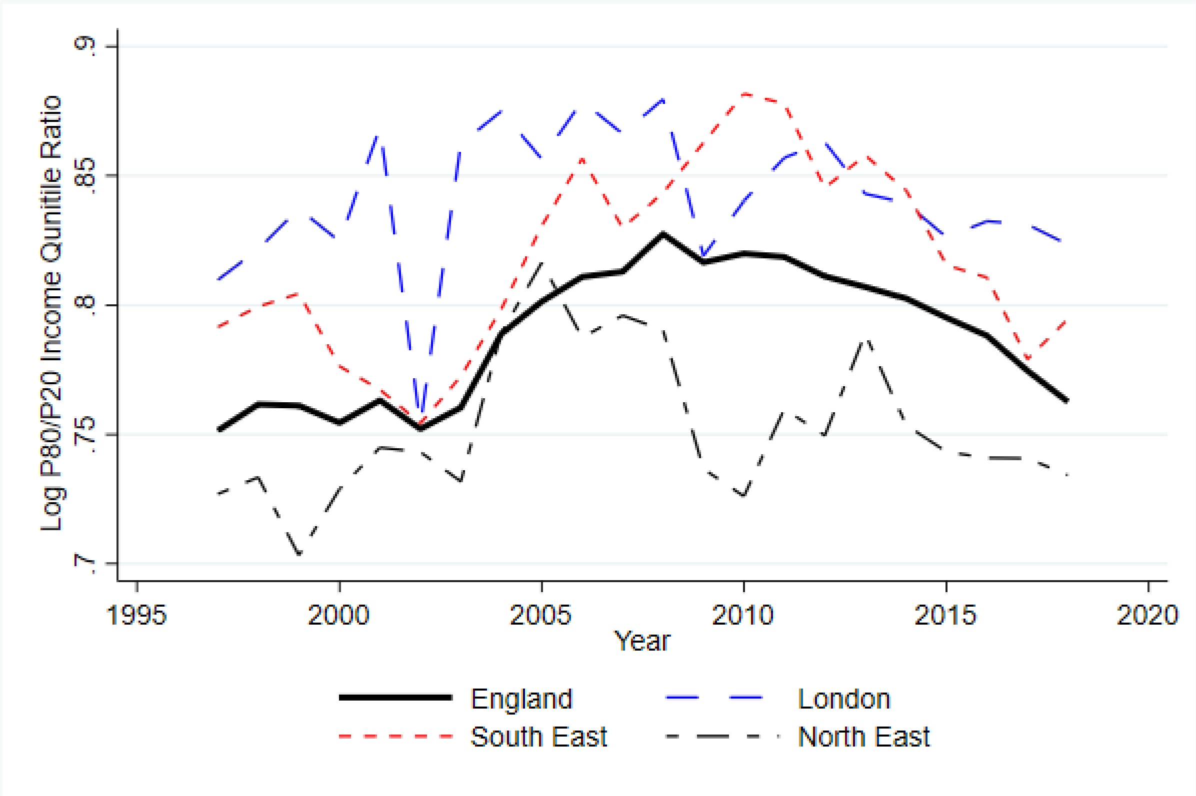
	(1)
	LLD w/o banking and real estate services
Log(adjusted local labor demand)	39.374***
	(10.617)
Av. refusal rate of major residential projects × log(adjusted local labor demand)	59.973***
	(8.786)
Share of developable land developed in 1990 × log(adjusted local labor demand)	79.100***
	(17.890)
Range between highest and lowest altitude × log(adjusted local labor demand)	22.675***
	(8.509)
Help to Buy (post-2015) x London dummy	-0.790
	(3.129)
LPA FEs	Yes
Year FEs	Yes
Observations	7,555
Number of LPAs	344
Kleibergen-Paap F	9.73

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The adjusted local labor demand measure is constructed from an index for the service sector excluding banking and real estate services (all other indices unchanged).

Appendix C: Appendix Figures

Appendix Figure C1

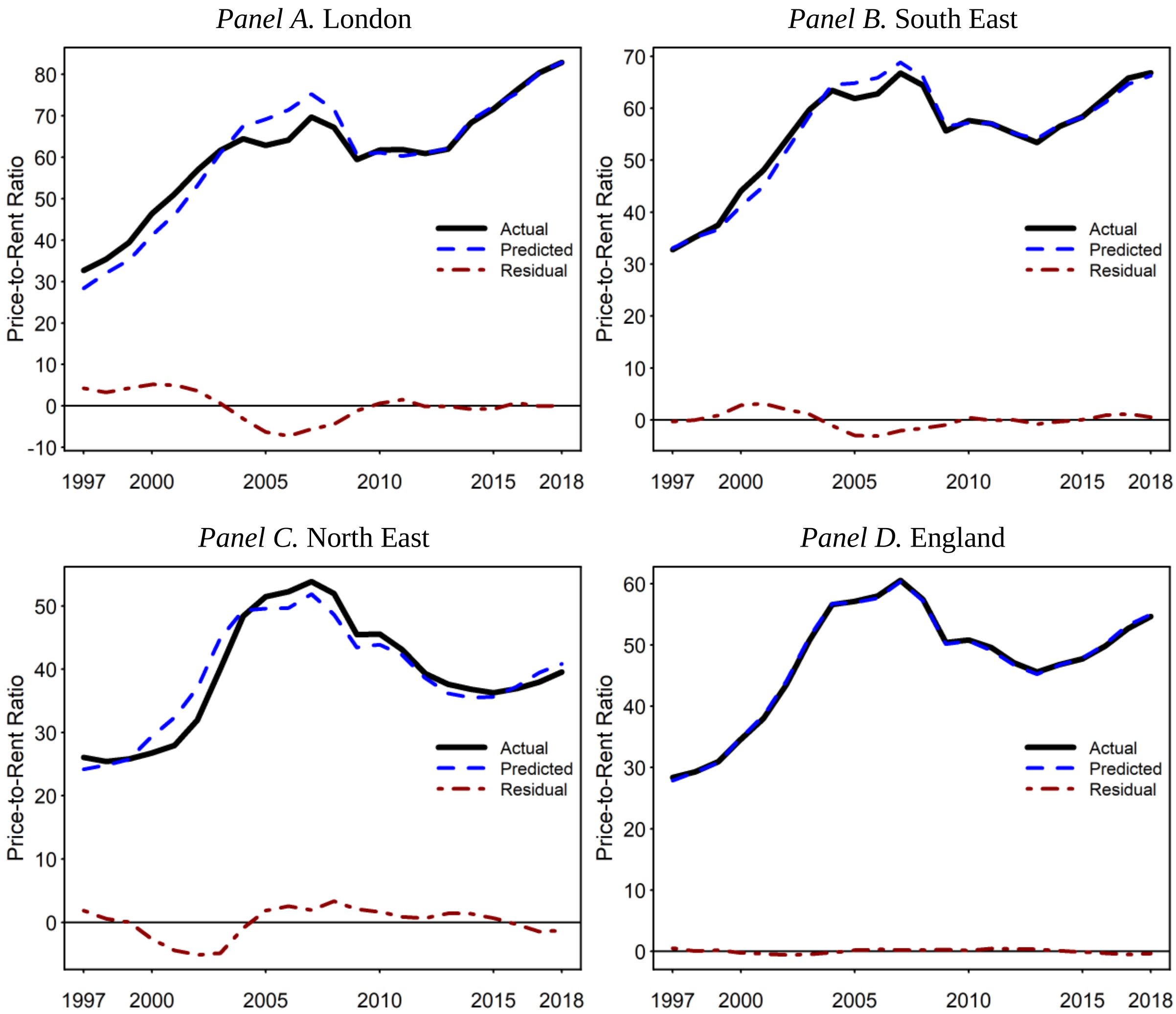
Income Inequality in England, London, the South East, and the North East of England, 1997-2018



Notes: The graph displays the average log ratio of the 80% income quantile to the 20% income quantile at LPA level, aggregated to England and the government office regions London, the South East, and the North East. The data source is the Annual Survey of Hours and Earnings, Table 7.1a - Weekly pay for full-time male workers at workplace.

Appendix Figure C2

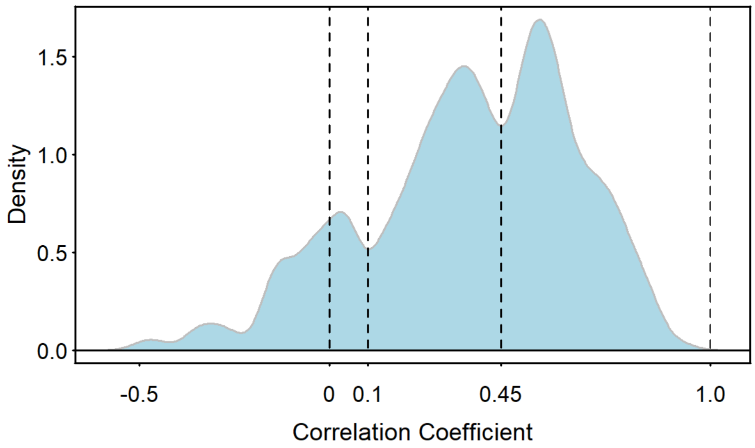
Actual and Predicted Price-to-Rent Ratios, and Residuals in London, the South East, the North East, and England



Notes: The graphs display the actual price-to-rent ratio for the respective region as a black solid line, along with the model-predicted value as a blue dashed line, and the residual as a red dotted-dashed line. Predictions are based on the models in Table 5, aggregated by year to the Government Office Regions (Panels A, B, C) and to England as a whole (Panel D).

Appendix Figure C3

Distribution of Correlation between Changes in Market Rents and PRP Rents at LPA-Level

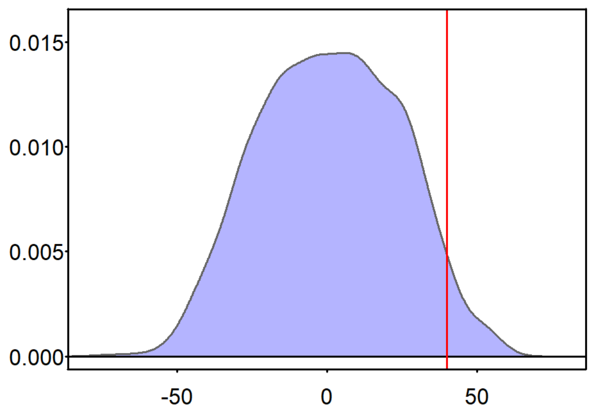


Notes: The graph displays the distribution of the correlation between changes in market rents and PRP rents at LPA-level. The dashed vertical lines indicate the three thresholds used to select the samples for Appendix Table B6 (correlation exceeding 0.0, 0.1, and 0.45).

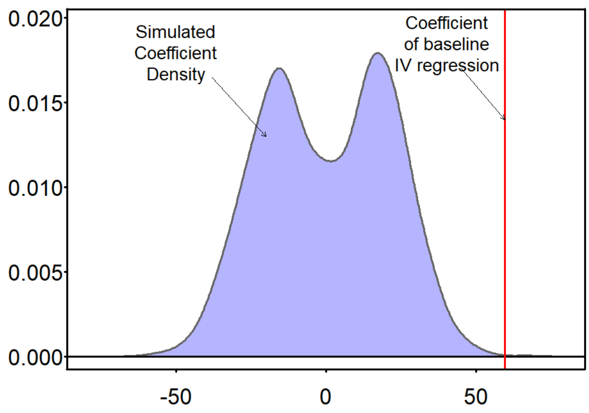
Appendix Figure C4

Placebo Test: Simulated Densities for the Baseline Regression Coefficients

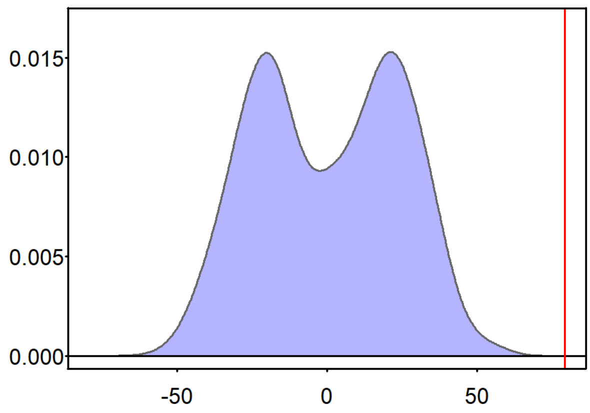
Panel A. Log LLD independent effect



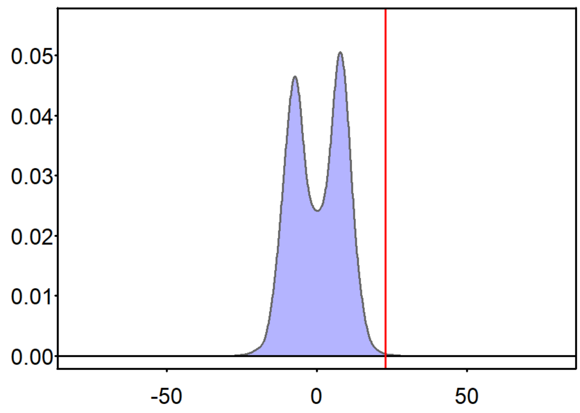
Panel B. Log LLD × avg. refusal rate



Panel C. Log LLD × share developed



Panel D. Log LLD × elevation range



Notes: All four graphs are based on the model displayed in Table 4. The graphs display the coefficient distributions from 2,000 simulated placebo labor demand measures, which are used instead of the shift-share labor demand measure. The red vertical bars indicate the locations of the baseline coefficient estimates.

Online Appendix – *Not for Publication*

Online Appendix O-A: Detailed Data Description

This online appendix provides details on the various sources and computation of variables used in our empirical analysis.

House prices. We extend and refine the house price panel of Hilber and Vermeulen (2016) from 2008 to 2018. We use the same composition adjustment to calculate average nominal house prices by LPA and year from the Price Paid Data of the UK Land Registry. The Price Paid Data contain all property sales in England of properties sold for full market value. The 1974 to 1994 panel is based on transactions recorded in the Survey of Mortgage Lenders. We drop transactions made under the Right-to-Buy scheme. The scheme allowed tenants in council housing to buy their housing units at a substantial discount. We append the full period for which the Price Paid Data are available, 1995 to 2018, to the adjusted 1974 to 1994 panel from Hilber and Vermeulen (2016). We deflate the nominal index by the RPIX.

Labor demand shock. We follow the methodology from Hilber and Vermeulen (2016). Specifically, we use industry shares at LPA-level from 1971 and Standard Industrial Classification (SIC) weights. We use seven broad industries.

The 1971 industry shares come from the Census of Population 1971. Like Hilber and Vermeulen (2016), we combine two national time series of employment growth by industry in order to arrive at a time series that covers the whole period, 1971 to 2018. The Census of Employment – Employee Analysis disaggregates employment of male fulltime employees in England into three-digit 1968 SIC categories. It is available from 1971 to 1978. Table O-A1 shows the disaggregation of employment for 1971 at the national level, for the Census of Employment and the Census of Population. Differences are attributable to the fact that unlike the Census of Population, the Census of Employment excludes women, part-time workers, and the self-employed.

Table O-A1

Industry Composition of Employment in 1971

Industry, as described in Census	% of total employment in 1971 England (Census)	 Great Britain (Employer Survey)
Agriculture	2%	2%
Mining	1%	3%
Manufacturing	35%	43%
Construction	7%	8%
Utilities; Transport	8%	12%
Distribution & Services	39%	24%
National & Local Government Service & Defence	7%	7%
Total	100%	100%

Source: Hilber and Vermeulen (2016).

We rely on weights proposed by Hilber and Vermeulen (2016) to deal with the various changes in the UK's industrial classification system. We use these weights to distribute industries from the more recent, finer classification systems to the classification system used in 1971.

For the period from 1978 until 2018, we use the Workforce Jobs by Industry data of employment by all fulltime workers in the UK, disaggregated to broad industries (one-digit 2007 SIC). The Office of National Statistics provides these data, drawing on employment and labor force surveys. Consistent with the 1971 Census of Population, this data includes the self-employed and women, but it excludes part-time workers.

The time series have one overlapping year, which allows us to calculate internally consistent growth rates. We use them to form industry-level employment indices for England as a whole, where 1971 is the base year. We then use the development of an industry's employment at the national level to extrapolate local employment in that industry in a given year, by simply multiplying the index value in that year with the industry's employment in the LPA in 1971. Our productivity shock measure is the sum over the extrapolated employment in all seven industries.

Share of greenbelt land in 1973. One of our instruments for the average refusal rate is the share of greenbelt land in 1973. In order to construct the variable, we digitized a map of recreational land in Great Britain (Lawrence 1973). The map provides information on greenbelts designated prior to 1973. We match the map with LPA delineations of 2001 and use geographic information software to calculate the share of designated greenbelt land in each LPA in 1973.

Market rents, 2010 to 2018. The rents data are taken from the "Private Rental Market Statistics" provided by the Valuation Office Agency. The Valuation Office Agency conducts surveys to collect data on rents. The Valuation Office Agency publishes average rents separately for different dwelling unit types (by number of rooms) for periods of 12 months (bi-annually, in March and October). We use the March publication and assign it to the same year. As an example, the March 2015 publication covers March 2015 to February 2016 and it was assigned to the year 2015 in the panel. We follow the same aggregation strategy as for the house price index. We first calculate the average share of each dwelling unit type by LPA and use these shares as aggregation weights in the second step. The nominal average rent by LPA and year is the weighted sum of mean rents reported for each category in that LPA and year. We deflate the nominal rents by the RPIX.

Private Registered Provider rents, 1997 to 2018. The uk.gov Table 704 of the UK Housing Statistics reports mean rents charged by Private Registered Providers (PRP), by year (1997 to 2018), and LPA. The statistic only includes larger PRPs with more than 1,000 beds and refers to self-contained units. PRP rents are subject to a rent ceiling that is pegged to the current market rent. We deflate the nominal rents by the RPIX. For more details on the definition of the rent ceiling, see the Guidance on Rents for Social Housing, Department for Communities and Local Government (now: Department for Levelling Up, Housing & Communities), May 2014, https://www.gov.uk/government/publications/guidance-on-rents-for-social-housing.)

Online Appendix O-B: Additional Tables

Table O-B1

Specifications separate for Periods with Positive and Negative Labor Demand Shocks – Results for Log Real House Prices and Log Real Rents

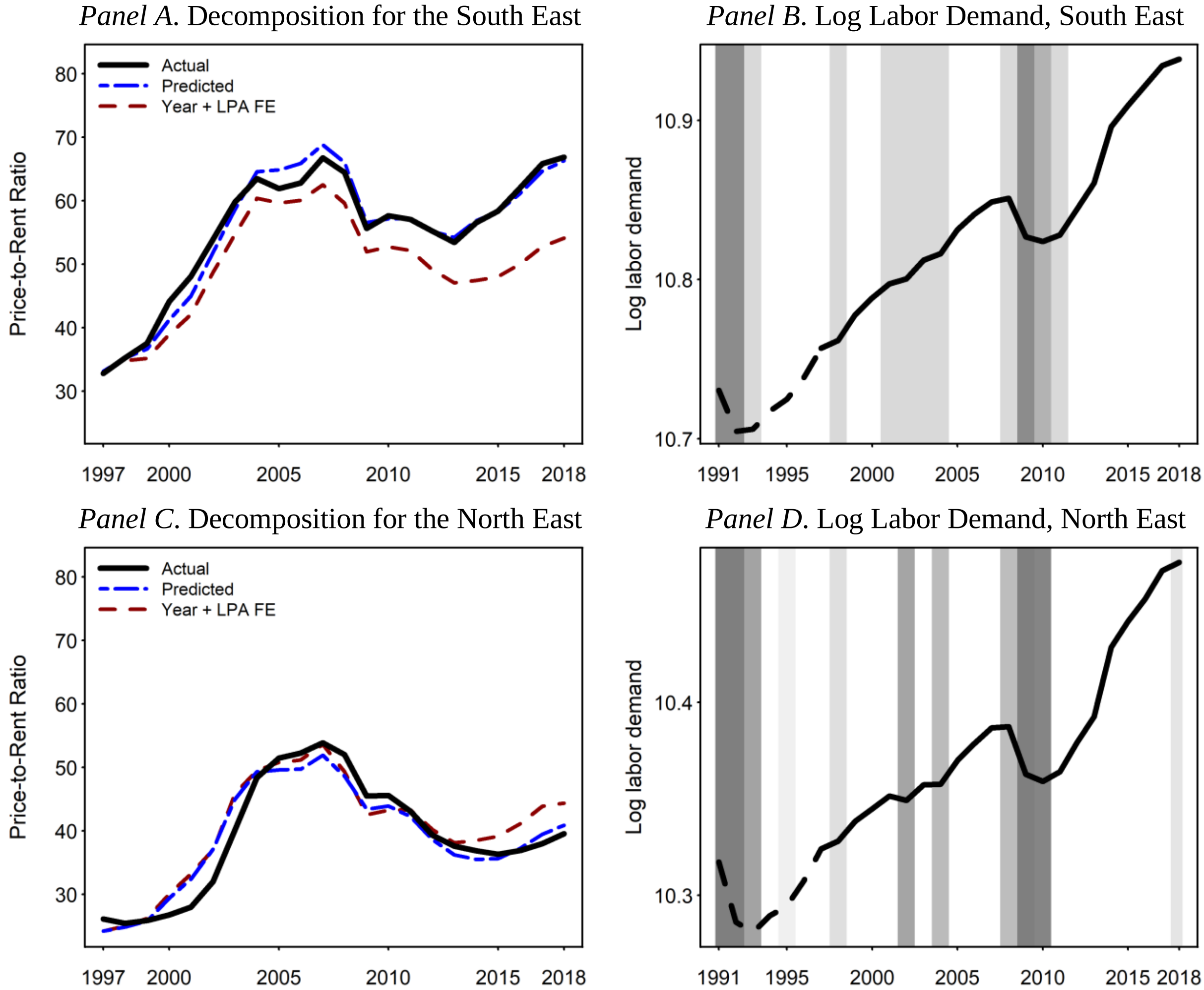
	(1)	(2)	(3)	(4)
	Prices 2SLS [a)] 1997-2018 [b), c)] △LLD>0	Prices 2SLS [a)] 1997-2018 [b), c)] △LLD≤0	Rents 2SLS [a)] 1997-2018 [c)] △LLD>0	Rents 2SLS [a)] 1997-2018 [c)] △LLD≤0
Log(local labor demand)	0.002	-0.467	-0.033	0.059
	(0.188)	(0.329)	(0.144)	(0.190)
Av. refusal rate ×	0.863***	0.077	0.309***	0.032
log(local labor demand)	(0.138)	(0.305)	(0.078)	(0.287)
Share developed ×	1.182***	-0.668*	0.527***	-0.176
log(local labor demand)	(0.277)	(0.361)	(0.090)	(0.140)
Altitude range ×	0.239*	-0.241***	0.144**	-0.097*
log(local labor demand)	(0.141)	(0.075)	(0.064)	(0.058)
Help to Buy (post-2015) ×	0.021		-0.050***	
London dummy	(0.049)		(0.016)	
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Observations	6,254	1,248	6,254	1,248
Number of LPAs	341	341	341	341
Kleibergen-Paap F	9.001	6.985	9.001	6.985

Notes: Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. [a)] First stage results are reported in Table B2. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, and population density in 1911 (persons per km²). [b)] Observations with missing rental data removed to make price and rent specifications comparable. [c)] LPAs w/o periods of decreasing local labor demand, as well as PRP vs. market rent outliers (mean log market rent > 7.5, based on Figure 4) removed to make the geographic extent of the sample (i.e., 341 LPAs) comparable.

Online Appendix O-C: Additional Figures

Figure O-C1

Decomposition of the Price-to-Rent Ratio and Log Labor Demand in the South East and the North East of England

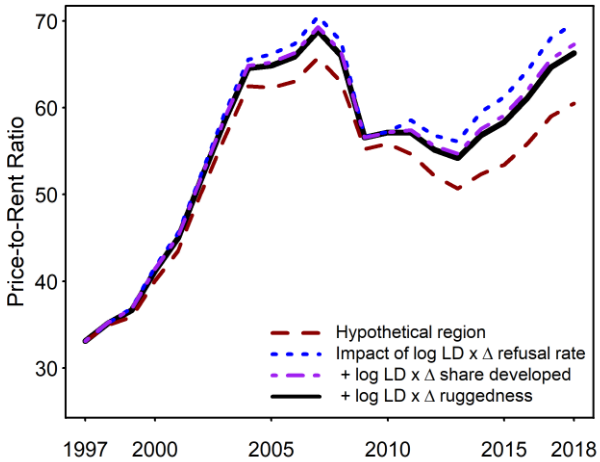


Notes: Panels A and C display the actual (solid black line) and predicted (dashed blue line) price-to-rent ratio, as well as the evolution of the price-to-rent ratio that is attributed to the fixed effects and Help-to-Buy (dashed red line), based on the models in Table 5, for the South East and the North East of England, respectively. The models from Table 5 were used to compute LPA-level predictions, that were aggregated to the Government Office Region of London, employing the number of households in each LPA in 2011 (Census) as weights. Panels B and D display the corresponding labor demand variable. In years with dark shading, all LPAs experienced decreasing demand. Medium-grey shading indicates periods with more than 50%, but less than 100% of LPAs experiencing decreasing demand. Light-grey shading represents periods with less than 50%, but more than 0% LPAs with decreasing demand. In periods without shading, demand increased in all LPAs.

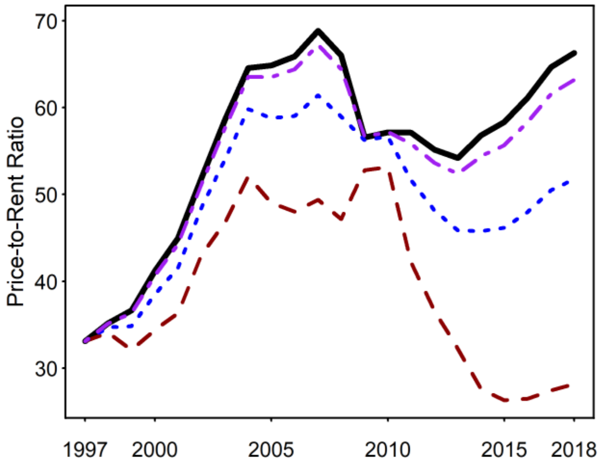
Figure O-C2

Counterfactual Decomposition Relative to a Location with All Supply Constraints at the 10% Sample Quantiles

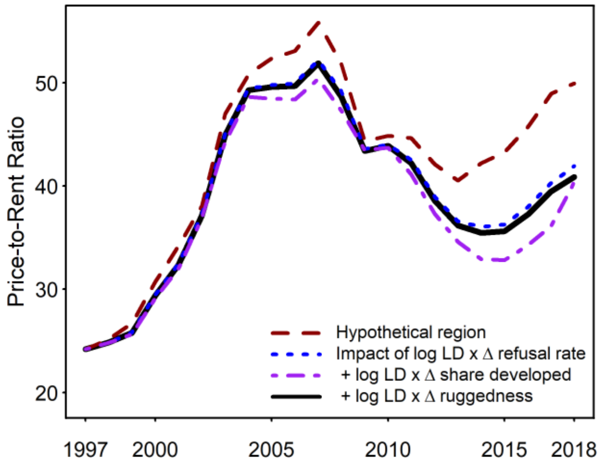
Panel A. Comparison of the South East to a Location with Average Supply Constraints



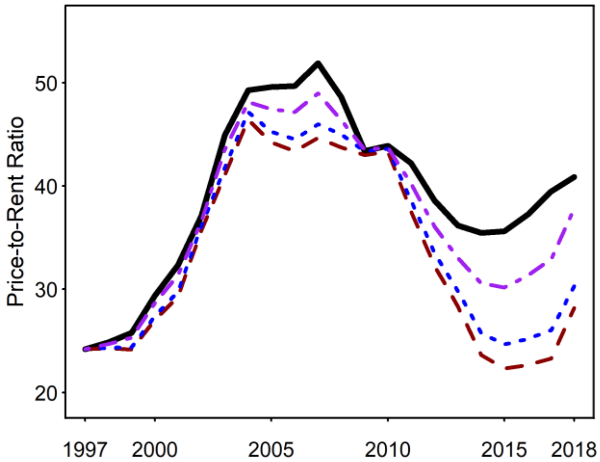
Panel B. Comparison of the South East to a Location with Lax Supply Constraints



Panel C. Comparison of the North East to a Location with Average Supply Constraints



Panel D. Comparison of the North East to a Location with Lax Supply Constraints



Notes: All graphs are based on the model displayed in Table 5. The model was used to compute LPA-level predictions, that were aggregated to Government Office Regions, employing the number of households in each LPA in 2011 (Census) as weights. Panel A compares the prediction for the South East (black solid line) to the prediction for a hypothetical location with average supply constraints (dark-red dashed line) and decomposes the difference into the impact of the local labor demand interactions with the differences in local regulatory restrictiveness (blue dotted line), the share developed land (purple dashed-dotted line), and ruggedness (the difference between the purple dashed-dotted line and the black solid line). Panel B repeats this exercise for a hypothetical location with all three supply constraints at the respective 10% quantile. Panels C and D show the same comparisons for the North East of England.

Online Appendix O-D: Idiosyncratic Risk Measures

In Section 3.6, we employ two time-varying measures of idiosyncratic investment risk at LPA level. We use data from the Land Registry to calculate LPA-year-level measures of idiosyncratic price risk. The first measure is based on Giacoletti (2021), who uses the concept of Local Market Equivalent, LME, defined as the abnormal performance of a house resale:

$$LME_{i,t_0,t_1} = \frac{\frac{P_{i,t_1}}{P_{i,t_0}}}{\frac{Q_{t_1}}{Q_{t_0}}} - 1 - \frac{D_{i,t_0,t_1}}{P_{i,t_0}}.$$

The numerator of the first term is one plus the capital gain of the individual house, where P_{i,t_0} and P_{i,t_1} are the sales prices of house i in years t_0 and t_1. The denominator is one plus the increase in local house prices between years t_0 and t_1, where Q_t is the local house price index. D_{t_0,t_1} are discounted maintenance expenditures. In Giacoletti (2021), this term does not turn out to be an important driver of idiosyncratic investment risk. Since we do not have detailed data on maintenance expenditures, we ignore this term.

Giacoletti (2021) then defines $lme_{i,t_0,t_1} = ln(1 + LME_{i,t_0,t_1})/\sqrt{t_1 - t_0}$ and regresses

$$lme_{i,t_0,t_1} = x'_{i,t_0}\beta + \psi_{postcode} + \phi_{t_1-t_0} + u_{i,t_0,t_1}.$$

Here, x'_{i,t_0} are controls including the initial purchase price P_{i,t_0}, $\psi_{postcode}$ is a postcode fixed effect, and $\phi_{t_1-t_0}$ is a holding period fixed effect. u_{i,t_0,t_1} is the residual. We run this regression separately for each LPA. Giacoletti (2021) interprets variation in u_{i,t_0,t_1} as capturing investment risk in period t_1, implicitly assuming that the initial purchase price P_{i,t_0} is exogenous, and that it captures the fundamental value of the property. The idiosyncratic volatility in year t in that LPA is given by the empirical standard deviation of $u_{i,t_0,t}\sqrt{t_1 - t_0}$, calculated over all observations with a repeated sale in period t.

We construct an alternative, simpler measure that accounts for the fact that u_{i,t_0,t_1} potentially depends on idiosyncratic volatility in both the initial purchase period t_0 and the sale period t_1. This measure is defined via the following FE-OLS regression at property level:

$$\ln P_{it} = \psi_i + \phi_t + x'_{it}\beta + u_{it}.$$

ψ_i and ϕ_t are property- and year-fixed-effects and x'_{it} are time-variant controls (in our case, this is a dummy for the property being newly constructed in period t, and a dummy for month of sale). We run this regression separately for each LPA.

Because ψ_i capture all time-invariant determinants of property i's price, ϕ_t is essentially a repeated-sales price index. We ignore unobserved changes to the property's characteristics, and changes in the valuation of the property's characteristics over time. Apart from this caveat, u_{it} captures the idiosyncratic component of the sales price in period t. We use the empirical standard deviation of u_{it}, calculated over all observations with a sale in period t as the idiosyncratic risk measure for year t in the LPA.