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

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


## Abstract

In most countries – particularly in supply constrained superstar cities – house prices have risen much more strongly than rents over the last two decades. We provide an explanation that does not rely on falling interest rates, changing credit conditions, unrealistic expectations, rising inequality, or global investor demand. Our model distinguishes between short- and long-run supply constraints and assumes housing demand shocks exhibit serial correlation. Employing panel data for England, our instrumental variable-fixed effect estimates suggest that in Greater London 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.



**JEL classification:** G12, R11, R21, R31, R52.

**Keywords:** House prices, housing rents, price-to-rent ratio, price and rent dynamics, housing supply, demand persistence, 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, 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 – but not in all – countries, house prices have risen much more rapidly over the past 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.

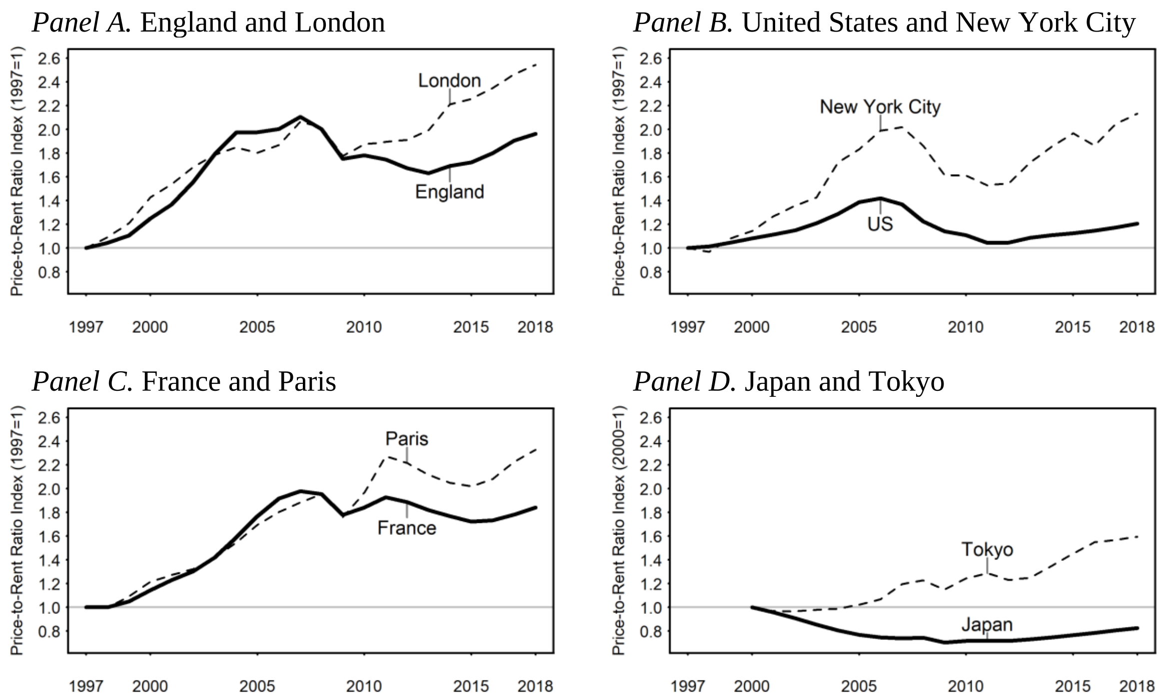
Figure 1 illustrates the stylized fact of a rising and cyclical price-to-rent ratio for England, France, and the United States. While in England the house price-to-rent ratio has almost doubled between 1997 and 2018 (our sample period), in France and the U.S. it has risen by 84% and 21%, respectively. This stylized fact is even more pronounced for the corresponding superstar cities. In London and Paris, the price-to-rent ratios have risen by a staggering 153% and 133%, respectively, between 1997 and 2018, while in New York City house prices have still grown more than twice as strongly as free-market rents (see the dashed lines in Figure 1). In all these cases the price-to-rent ratio has risen 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. Here the price-to-rent ratio has been falling over the last 20 years, despite a decrease in the real rate of interest.[1] However, in Tokyo, where population has been growing, the price-to-rent ratio increased by 60%.

More generally, as Figure 2 portrays for England, the increase of the price-to-rent ratio since 1997 varies enormously across regions. Whereas in Greater London[2] and the South East the increase in the price-to-rent ratio has been above the country's average, in the North East it has been significantly below average (Panel A). As illustrated in Panel B, the price-to-rent ratios have been higher in London and the South East than in the North East throughout our sample period.[3] However, the regional differences widened remarkably during the two boom periods 1997-2006 and post-2009, while they narrowed during the Great Financial Crisis.

[1] According to the World Bank, Japan's real interest rate declined from 3.5% in 2000 to 1.1% in 2017.
[2] When we refer to 'Greater London' or 'London' we mean the Greater London Authority.
[3] There are several plausible explanations for this stylized fact including spatial differences in (i) the risk premia (Amaral *et al.* 2021) and (ii) the depreciation rate—see Section 3.1. Explaining these persistent cross-sectional gaps is not the focus of our paper.

Figure 1

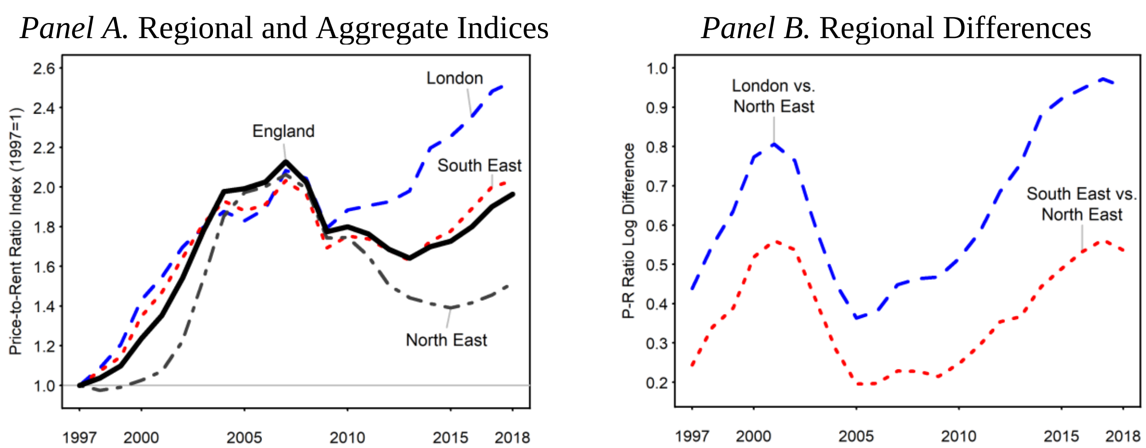
*Price-to-Rent Ratio Indices for Selected Countries and Superstar Cities (1997-2018)*




*Notes:* The series for England and London are based on transaction prices (Land Registry) and Private Registered Provider rents (MHCLG 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

*Price-to-Rent Ratio Indices in England*




*Notes:* Panel A 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. Panel B displays log differences between the price-to-rent ratios for London vs. the North East and for the South East vs. the North East. House prices are based on transactions (Land Registry). Rents are Private Registered Provider rents taken from MHCLG 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 rate of interest or with unprecedented availability of housing credit may explain a significant fraction of the price-to-rent dynamics at national level, macroeconomic conditions alone cannot account for the systematic differences 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 much more pronounced in economically thriving and tightly supply constrained superstar cities, even when holding macroeconomic conditions constant. We show that the stylized facts are consistent with a simple model that distinguishes between local short- and long-run supply constraints and assumes – consistent with our data – that local housing demand shocks exhibit serial correlation.

Agents in our simple two-period model understand that housing demand shocks are serially correlated, but they do not have perfect foresight. A given housing demand shock triggers an immediate – short run – reaction of supply. Agents then adjust their price and rent expectations, which in turn depend on expected future housing demand shocks and the expected response of housing supply in the long run. In this setup, (i) the price-to-rent ratio increases in response to a positive shock if local housing supply is sufficiently constrained, (ii) the more so, the more supply constrained the location, and (iii) this interaction effect is amplified during periods with strong demand persistence. Moreover, provided the housing supply curve is inelastic (kinked) 'downwards', (iv) the price-to-rent ratio decreases in response to a negative housing demand shock, irrespective of the 'upward' supply price elasticity.

In our empirical analysis, we work out the impact of the interaction between local housing demand and local housing supply constraints as well as the dynamics of price- and rent-adjustments over the local business cycle. 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 housing demand. The latter is an important aspect, as housing 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)[4] and annual data from 1974 to 2018 (for house prices) and 1997 to 2018 (for rents). 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

[4] LPAs are the local authorities (or councils) that are responsible for the execution of planning policy. In this sense, 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 January 9, 2020.

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 that (i) the impact of local labor demand shocks on the price-to-rent ratio depends strongly on local supply constraints and (ii) this interaction effect is amplified during business cycle-periods with high persistence in demand shocks, consistent with theory, thus lending support to our novel proposed mechanism.

In Greater London, where supply is seriously constrained, local labor demand shocks in conjunction with local supply constraints explain 63% of the increase in the price-to-rent ratio since 1997. The year fixed effects in our panel fixed effects analysis explain the remaining 37%. 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 can explain the bulk (84%) 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 capture 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 fixed effects also capture the impact of persistent common housing 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.

Consistent with the supply curve being kinked, the stronger increase in the price-to-rent ratio in Greater London during our sample period is entirely driven by *positive* local labor demand shocks and their interactions with supply constraints. *Negative* shocks have a negative effect on the price-to-rent ratio, but, consistent with theory, this negative effect is independent of local supply constraints.

We also provide evidence discounting the possibility that our key findings are driven by alternative mechanisms (irrational exuberance, segmented markets in conjunction with rising income inequality, changing mortgage finance conditions, impact of Help to Buy, rent stickiness in existing contracts, or global investor demand for second homes in superstar cities).

Our paper 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, especially 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) suggested 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 of the literature emphasizes the demand side and the financing of housing. 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. Much of this literature again focuses on the United States. Himmelberg *et al.* (2005) suggest that it was easily available credit in the years preceding the Great Financial Crisis that led to low interest rates, which in turn boosted housing demand and house prices.[6] 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 and demonstrate that these can easily be reconciled 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 occurred 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.

Recent work in the UK has instead focused on the sustained decline in real interest rates over the last two decades and the tightening of credit conditions in 2008. Miles and Monro (2019) rely on a user cost model to rationalize the increase in real house prices in the UK at macro-level. Their model-based predicted increase in real prices is driven almost exclusively by the unexpected fall in real interest rates and increases in real incomes between 1985 and 2018, with

[6] Other studies however question the importance of falling real interest rates in explaining the house price boom preceding the Great Financial Crisis. Kaplan *et al.* (2020) identify a shift in beliefs about future housing demand as the dominant force driving the boom-bust cycle. 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.

both components being equally important. Their model matches the observed increase in real prices between 1985 and 2018 but does not work as well for sub-periods. Moreover, for conceptual reasons, their analysis cannot inform about the relative importance of supply price elasticities for these relationships.

While the two strands of the literature have had little overlap, more recently, proponents in England (most prominently, Mulheirn 2019), the United States,[7] and elsewhere have pointed to the rising price-to-rent ratio as 'direct evidence' that "the housing shortage hypothesis [driven by a dysfunctional planning system] is misplaced". Moreover, rising global investor demand for parts of London (Badarinza and Ramadorai 2018) is invoked to justify the stronger increase in house prices and the price-to-rent ratio in the capital.

Our study reconciles the two 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 locations like London. Our study stresses the importance of local demand and supply side determinants especially in tightly constrained locations, alongside macroeconomic factors.

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 Molloy *et al.* (2020), Büchler *et al.* (2021) and Kaplan *et al.* (2020). Molloy *et al.* (2020) study the relationship between long-differences in prices and rents, and time-constant constraints to the supply of housing, finding a relatively stronger association between price changes and supply constraints.[8] Their theoretical explanation assumes a positive, constant growth rate of aggregate housing demand in a two-region dynamic setting. In such a setting, as long as the rate of new housing supply is sufficiently constrained in one region (relative to the other), housing supply in that region can never catch up with the change in demand, resulting in the expectation that future housing rents always exceed today's rents. Büchler *et al.* (2021) also study long-differences in prices and rents, during a period of rising housing demand. They focus on differences in local housing supply elasticities between locations, finding relatively larger elasticities for prices than for rents, as well as strong spatial differences related to supply constraints. The authors argue that prices react more strongly to demand shocks than rents because shocks lead investors to update their expectations of local risk premiums and rent growth rates, with the degree of updating depending on the share of sophisticated investors at a location.[9] Kaplan *et al.* (2020) propose a

[7] See Been *et al.* (2019) who critically assess the 'supply skepticism' arguments in the United States.

[8] Molloy *et al.* (2020) regress price changes on local housing supply constraints and covariates. This contrasts our approach of regressing price changes on the *interaction* of local supply constraints and changes in housing demand plus covariates. The latter allows us to consider the fact that local supply constraints may have a differential effect on house prices depending on the extent of local demand shocks.

[9] Our theoretical and empirical setups differ from Molloy *et al.* (2020) and Büchler *et al.* (2021) in important ways, leading to significant differences in the interpretation of the observed stylized facts. In particular, in contrast to the two other papers, our theoretical and empirical setups consider both positive and negative housing demand shocks, allowing us to investigate whether these shocks have symmetric effects. We find asymmetric effects that depend on local supply constraints. Consistent with our model, we find that the price-to-rent ratio increases in response to a positive shock only in locations with sufficiently constrained housing supply. Moreover, the price-to-rent ratio decreases in response to a negative housing demand shock, irrespective of the upward supply price elasticity. In contrast to Molloy *et al.* (2020), we consider agents who do not have perfect foresight and we allow housing supply to eventually catch up to local demand. In contrast to Büchler *et al.* (2021), we do not rely on

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, however, not explore why beliefs shifted. We shed light on this: In our setting rational agents form beliefs over the (local) business cycle. In addition, we explore the role of differences in local supply price elasticities and the causes of the cyclical spatial differences in price-to-rent ratios that we observe in the data.

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 and robustness checks. In Section 4, we investigate the quantitative importance of the mechanism and explore alternative explanations. The final section concludes.

## 2 Theory

To explain why not only house prices and rents but also the price-to-rent ratio respond more strongly to labor demand shocks when housing supply is tightly constrained, we develop a simple model of local housing markets that differ in their short- and long-run housing supply elasticities. The mechanism we propose builds on two crucial assumptions: (1) short-run housing supply is less price elastic than long-run supply because of binding short-run planning and construction lags, and (2) local housing demand shocks exhibit serial correlation, which is a feature of our data.

Moreover, we assume that locations with tight long-run housing supply constraints also face more severe short-run planning and construction lags. There are several reasons for this: First, the delay rate of planning applications increases with 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 the development. Third, it is more difficult to build in locations that are more rugged, which arguably increases construction time. For all these reasons, short- and long-run elasticities are highly likely positively correlated.

Since 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 rents.[10] As long as supply is not perfectly inelastic in the short-run, 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 2), future expected rents may be higher despite the larger long-run supply elasticity. In that case, prices react more strongly to an initial demand shock than rents. This implies that price-to-rent ratios increase in reaction to (strongly) serially correlated positive housing demand shocks, and this increase can

exogenous differences in investor beliefs across locations. In our case, our findings are consistent with agents following the *same* rule about updating expectations in all locations.

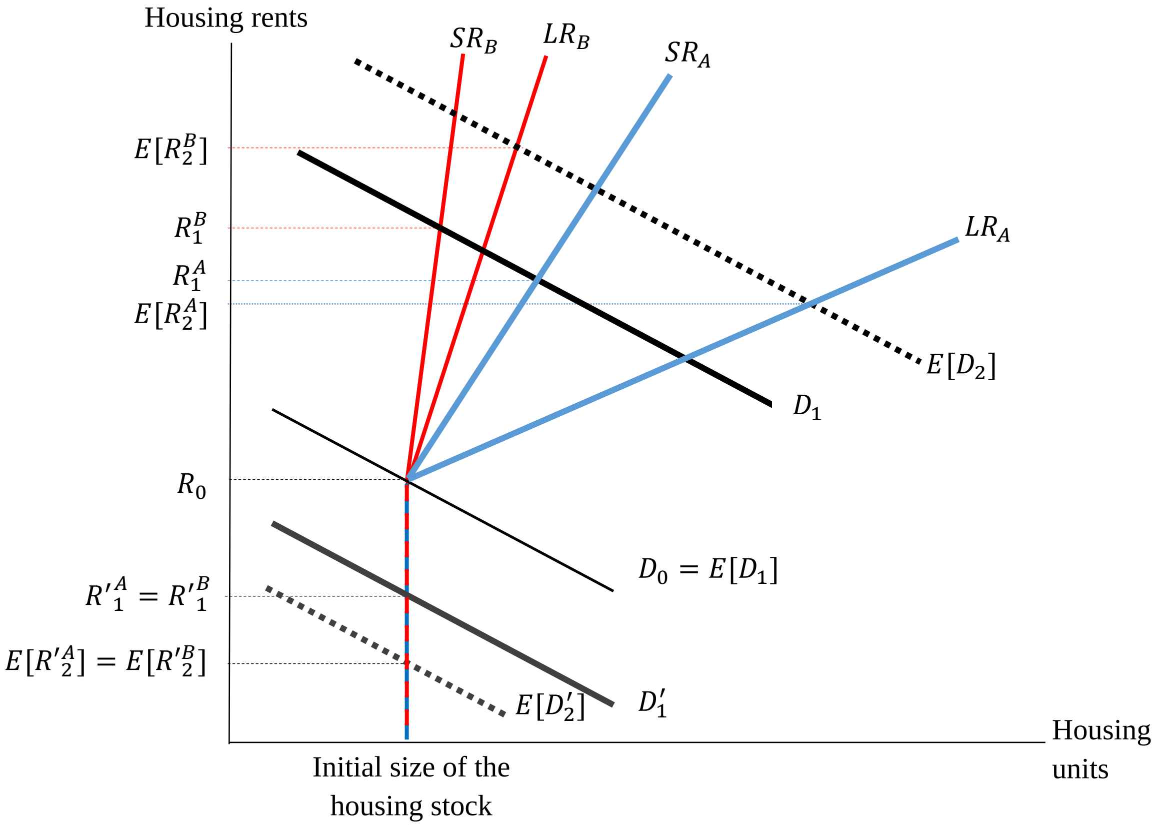
[10] We abstract here from the possibility of rent control and sticky rents. Private rents in England are not subject to rent control, and landlords can adjust rents freely during a tenancy.

be expected to be stronger in locations with more inelastic long-run supply constraints, as the latter attenuate the long-run supply response.

Figure 3 provides the intuition for these predictions. In location A, the upper parts of the housing supply schedules for short- and long-run housing supply are less steep 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), which shifts the demand schedule from $D_0$ to $D_1$, increases rents (and prices) up to the intersections with the 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 $E[D_2]$ and 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.

In equilibrium in period 0, $E[R_1] = R_0$, hence $P_0 = R_0(1 + 1/(1 + r))$. The price in period 1 is $P_1 = R_1 + E[R_2]/(1 + r)$. 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 (see Figure 3).

Figure 3
*Theoretical Mechanism*



We now turn to the model. We start with a 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 fundamentally 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's 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), before discussing the (expected) long-run equilibrium outcome (period 2 captures the long-run). Panel B of Figure 2 suggests that the time difference between the short- and long-run periods may well be five to ten years. This setting has the advantage of being simple while still maintaining the key mechanism we have in mind.[11]

Assume that the location is characterized by a short- and a long-run housing supply elasticity, which we take to be exogenous[12], as well as by location fundamentals $a_t$ (amenities) and $\omega_t$ (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 agents in the model are renters. Investor-landlords willing to pay the present discounted value of the housing unit determine the price of housing.[13]

We assume the location is in an equilibrium, that is, the expected demand shock in period 0 is zero. Households are indexed by $i$. They have an outside option that yields utility $\bar{u}$, which we normalize to $\bar{u} = 0$. Their utility from living in the location in a given period is $w_t - R_t - \eta_i$, whereby $R_t$ is the rent in period $t$ and households have an idiosyncratic (dis-)taste for the representative location, represented by a parameter $\eta_i \sim \mathcal{U}_{[0,\phi]}$. Here, $\phi$ is a taste dispersion parameter. If $\phi$ is small, households have a relatively stronger taste for the location, on average.

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

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

[11] 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 construction capacity limit (per period), as in Wheaton (1999).

[12] 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.

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

The resulting initial equilibrium rent level in period 0 is $R_0 = w - \phi S_0$. We assume that a shock $\varepsilon$ to local wages in period 1 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 short-run housing supply function is given by

$$S_1 = S_0 + \delta\beta(R_1 - R_0). \tag{2}$$

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 $\delta$ means that short-run supply is less elastic relative to long-run supply of the location. $\beta$ captures the location's long-run supply elasticity. Hence, a smaller $\beta$ 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. This connection of the short- and long-run supply price elasticities capture the idea that short-run planning and construction lags are related through several features of the regulatory environment, as well as through the geographical and topographical constraints to housing supply.

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

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

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 $\phi$ is small).

After having observed the demand shock $\varepsilon$, the long-run expected demand is $E[D_2] = (w + \varepsilon(1+\gamma) - E[R_2])/\phi$, where $E[R_2]$ is the expected long-run rent level. The long-run supply curve is $S_2 = S_0 + \beta(E[R_2] - R_0)$, which 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 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 $\beta$ or $\phi$ are small). Therefore, similar relationships hold for the house price in period 1, which we define as $P_1 = R_1 + E[R_2]/(1+\mathrm{r})$. Here, $r$ is the discount rate that is exogenous to the model. 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 $\beta$ is small, which reduces the long-run supply response, or if local housing demand is relatively elastic (i.e., when $\phi$ is small).[14] In that case, the earnings 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 can summarize these results as follows:

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.

*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

[14] 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).

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 the effects predicted by the model are transitory. The example of the sustained population decline in Japan – causing the price-to-rent ratio to fall – and the simultaneous population growth in Tokyo – triggering an increase in the price-to-rent ratio – shows that these transitions can last for decades.[15]

According to *Proposition 3*, differential positive rent growth expectations emerge during the boom phase of a local business cycle because supply constraints amplify the impact of positive housing demand growth expectations and persistence of demand shocks may be especially strong in the midst of a boom (see below for supporting evidence on the latter). As Figures 1 and 2 show, these local upswings can last a decade (and possibly longer). Rent growth expectations revert to zero 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 1974-2018 for house prices and 1997-2018 for rents. Summary statistics of the key variables are reported in Table 1.[16]

*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. 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

[15] Our model considers local business cycles and local supply constraints. However, the mechanism of our model in principle also applies to the macro-level. 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).

[16] We provide more detail and background information in Online Appendix O-A.

indices at LPA-level. We refine the index by dropping 'Right to Buy' transactions[17] from the Council of Mortgage Lenders data and deflate the nominal indices by the national-level retail price index net of mortgage payments (RPIX). The full house price series covers the period from 1974 to 2018.

Table 1
*Summary Statistics*

| | Mean | Standard Deviation | | | Min. | Max. |
|---|---|---|---|---|---|---|
| | | Overall | Between | Within | | |
| **A. 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. 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 |
| Log(local labor demand) [b)] | 10.8 | 0.64 | 0.64 | 0.05 | 8.35 | 13.16 |
| **C. 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 |
| Log(local labor demand) [b)] | 10.8 | 0.62 | 0.62 | 0.05 | 9.26 | 13.16 |
| **D. 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 |

*Notes:* [a)] Based on house price transaction data. [b)] Log predicted employment, based on 1971 local industry composition and national employment growth.

We employ two measures for local rents. The first is based on Private Registered Provider (PRP) rents provided by the Ministry of Housing, Communities and Local Government (MHCLG), which are available from 1997 to 2018.[18] While some PRPs are for-profit

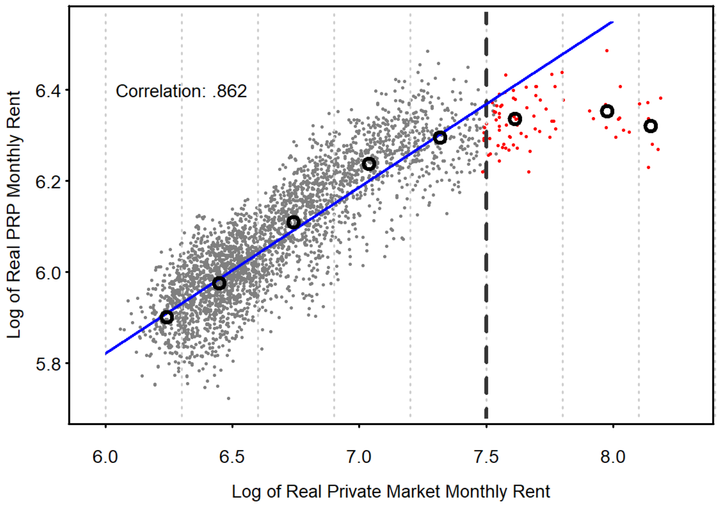
[17] The 'Right to Buy' scheme, implemented in 1980, permitted tenants in Council Housing to buy their homes at a discount that could be as high as 40% of the market value of the unit.
[18] 1997 is the first year with any available rental data for England at local level, see the gov.uk Live Table 704.

organizations, others are not-for profit. In all cases however, they have an incentive to maximize their rental income, subject to constraints; the latter group to be able to reinvest surplus income 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. The second measure uses mean private market rents, provided by the Valuation Office Agency. Private market rents are available from 2010 to 2018 and we construct a mix-adjusted index that holds constant the average dwelling size (number of rooms). We again deflate the rent measures by using the RPIX.

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. Our main analysis uses PRP rents because this allows us to cover a period of 22 years, with several (local) booms and busts. We use a simple rule based on a visual inspection of Figure 4 to deal with the potential discrepancies between PRP and market rents. That is, we exclude all LPAs with a mean log market rent exceeding 7.5.[19]

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 market rents in these LPAs exceeded 7.5. Market rents are available only starting in 2010 from the Private Rental Market Statistics collected by the Valuation Office Agency.

We measure the local price-to-rent ratio as the ratio of average house prices to average rents. The focus of the analysis is on explaining differential changes in the price-to-rent ratio between

[19] Our empirical findings are not sensitive to using alternative and more sophisticated rules. That is, in several robustness checks, we base our sample selection on the correlation between yearly changes in log PRP and log market rents and select LPAs where this correlation exceeds different thresholds. We also run regressions based on the full sample and with the private market rents as dependent variable and our main findings remain qualitatively unaltered.

locations, so that concerns relating to the comparability of housing units between the rent and price data are arguably less relevant.

*Implied Rental Return*

The mean price-to-rent ratio based on the PRP rents in all LPAs in 2018 was 56. In London it was 82. To make sense of these numbers, consider the inverse, which has an interpretation as the implied gross rental return (1/56 = 1.8% and 1/82 = 1.2%). From a finance perspective, in equilibrium, the gross rental return net of maintenance costs plus the expected rent growth ought to be equal to the return on a risk-free asset plus the risk premium for the investment in housing. Since the PRP rents are subject to a rent cap at 80% of market value, the implied market returns are 1.8%/0.8 = 2.25% and 1.2%/0.8 = 1.5%, respectively.

In 2018, the (risk-free) Rate of the Bank of England was 0.5%. Maintenance costs depend on the land value share – what matters for maintenance is the depreciation of the structure, not the value of the land. Under the assumption that the replacement costs of the structure are equal across locations, the price for a (new) home is the sum of a constant plus the location-specific land value. Based on appraisal estimates for land values in 2017 by the Valuation Office Agency, the average land value in England (London) is 6.0 (35.5) times higher than that in the North East, whereas house prices are only 1.9 (3.3) times higher. These numbers imply land value shares of 57% in England, 18% in the North East, and 89% in London.[20] Assuming an annual depreciation rate of the structure of 2%, the maintenance costs are thus 0.86% p.a. for England and 0.22% p.a. for London. If we further assume a risk premium of housing of 1.5%, we get implied expected real rent growth rates of about 0.6% p.a. for England and 0.7% p.a. for London, matching fairly closely the observed average real rent growth rates over our sample period of 0.5% for England and 1.0% for London. These numbers suggest that differential rent growth expectations (driven in part by differential supply constraints and in part by differential demand shocks) are a sensible candidate explanation for the spatial differences in implied rental returns – as suggested by our theoretical framework.

*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 MHCLG. The ‘refusal rate’ is simply the number of refused applications divided by the total number of applications in a given year. The refusal rate of ‘major applications’ (i.e., applications of projects consisting of ten or more dwellings) 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.

[20] With replacement costs c and land value x for the North East, we have $(c + x) / (c + 6.0 \times x) = 1.9$ when comparing the North East to England, and $(c + x) / (c + 35.5 \times x) = 3.3$ when comparing the North East to London, implying $x = 0.33$ and $c = 1.48$. The resulting value shares of the structure in the North East, England, and London are 0.82, 0.43, and 0.11, respectively.

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 local housing demand shifter is a Bartik (1991) shift-share measure that captures shocks to local labor demand. We follow Hilber and Vermeulen (2016) and use the Census 1971, which provides employment by industry for seven industries at LPA-level. National level employment growth by industry is derived from the Census of Employment (1971-1978) and the Office for National Statistics (1979-2018).[21] We use this data to predict local labor demand, $LLD_{it}$, in each LPA $i$ and year $t$ during our sample period.

Our theoretical model assumes that shocks to local housing demand exhibit a degree of 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. Panel A of 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, almost 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.[22]

We would however expect that the demand shock persistence varies strongly within a location over the local business cycle. That is, in the midst of a boom, it is arguably more likely that a positive shock is followed by further positive shocks. The same argument applies to a local bust. To test this hypothesis, we define five phases of the local business cycle, based on the deviation of $LLD_{it}$ from its local trend path (the latter is captured by a quadratic polynomial at LPA-level).[23] We consider changes in these residuals for the definition of the five cycle phases.

We define the first phase – the 'beginning of the boom' – as the first two periods (years) with consecutive positive changes. The second phase – the 'boom' – starts with the third consecutive positive change and ends when the number of consecutive positive changes exceeds the LPA's average by at least two periods. At this point, we define the boom as being 'overly long'. In very long booms, irrational exuberance may manifest itself, making results for this phase particularly interesting. This third phase ends with the first negative deviation from the trend. For the bust, we consider two phases ('beginning of the bust', 'bust'), disregarding 'overly long busts', because busts are typically shorter than booms (leading to small sample issues).

To assess persistence in each of these five phases, we then estimate the following regression

[21] We rely on weights proposed by Hilber and Vermeulen (2016) in order 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.

[22] The median degree of persistence hardly differs between locations inside and outside of the Greater London area (0.547 and 0.542).

[23] This fits the data better than a linear approximation.

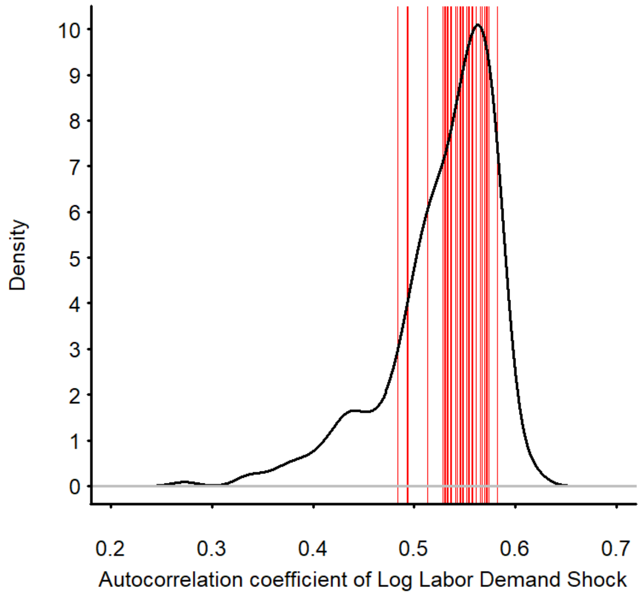
$$\Delta LLD_{it} = g_i + \sum_{k=1}^{5} \gamma^k \times \left(\Delta LLD_{i,t-1} \times \chi_{i,t-1}^k\right) + \eta_{it}. \tag{6}$$

Here, $g_i$ captures the local trend in LPA $i$, $\chi_{i,t-1}^k$ is an indicator for phase $k$, and the $\gamma^k$'s capture persistence in the different phases. Panel B of Figure 5 displays the estimated $\gamma^k$'s. The graph clearly shows that persistence is weaker at the beginning and towards the end of a boom, as well as at the start of a bust. It is by far strongest in the midst of a boom. Moreover, the variation over time is very substantially larger than the typical variation across space, as depicted in Panel A of Figure 5.

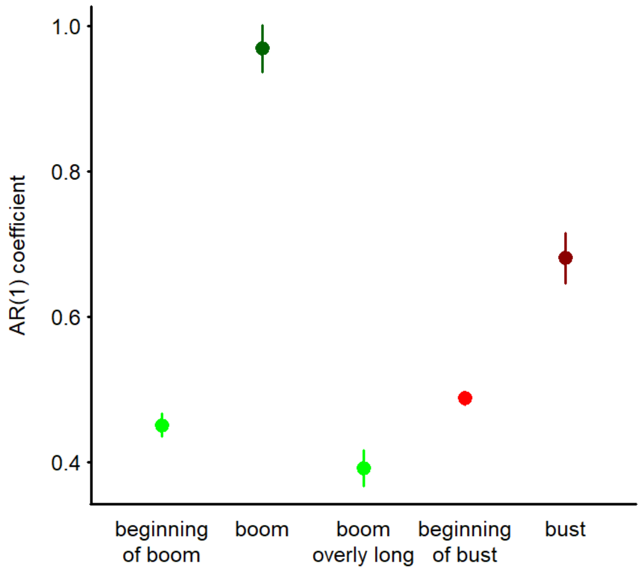
Figure 5

*Persistence in Local Labor Demand Shocks*

*Panel A.* Spatial Distribution of Autocorrelation Coefficients




*Panel B.* Persistence over the Local Business Cycle




*Notes:* The graph plots the autocorrelation coefficients of the local labor demand shock. *Panel A* shows the spatial distribution of persistence, with vertical red bars indicating LPAs in London. *Panel B* displays the coefficients by phase of the local business cycle (accounting for LPA-specific trends). The vertical bars indicate 95% confidence intervals (clustered by LPA).

### 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.

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.[24] 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,[25] 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.[26] 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 that restrictive LPAs – to meet their delay rate target – responded to the reform by delaying less and refusing more.[27]

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.

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

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 one or two of the three instruments.

*Identifying the Share of Developed Land*

The share of developable land developed in 1990 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 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

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

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 capture the 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: $\theta_0$ captures the impact of a labor demand shock on the outcome in an LPA with average supply constraints in all three dimensions. The coefficients $\theta_1, \theta_2$, and $\theta_3$ 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 and for different subsamples capturing the five phases over the local business cycle defined in Section 3.1 ('beginning of boom', 'boom', 'overly long boom', 'beginning of bust', 'bust') as well as 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)<br>OLS<br>1974-2018 | Log(Prices)<br>2SLS [a)]<br>1974-2018 | Log(Prices)<br>2SLS [a)]<br>1997-2018 [b), c)] | Log(Rents)<br>2SLS [a)]<br>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 | 0.188*** | 0.652*** | 0.863*** | 0.283*** |
| residential projects × log(LLD) | (0.069) | (0.118) | (0.123) | (0.071) |
| Share of developable land | 0.438*** | 1.099*** | 1.110*** | 0.504*** |
| developed in 1990 × log(LLD) | (0.148) | (0.117) | (0.253) | (0.083) |
| Range between highest and | -0.044 | 0.326*** | 0.203* | 0.124** |
| lowest altitude × log(LLD) | (0.041) | (0.108) | (0.122) | (0.056) |
| Help to Buy (post-2015) x | 0.242*** | 0.047* | 0.035 | -0.049*** |
| London dummy | (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 3. 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 km$^2$). [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 3. (All subsequent first stage results corresponding to Table 2 are also reported in Table 3.) 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.

Table 3

*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<br>(0.060) | 0.138**<br>(0.062) | 0.098<br>(0.075) | 0.198**<br>(0.080) |
| Altitude range × log(LLD) | -0.067<br>(0.052) | -0.392***<br>(0.041) | -0.066<br>(0.046) | -0.336***<br>(0.034) |
| Change in delay rate × log(LLD) | -0.080*<br>(0.043) | 0.017<br>(0.046) | -0.082*<br>(0.042) | -0.014<br>(0.039) |
| Share Labour vote in 1983 × log(LLD) | -0.512***<br>(0.070) | 0.245***<br>(0.050) | -0.588***<br>(0.041) | 0.277***<br>(0.041) |
| Share greenbelt in 1973 × log(LLD) | 0.289***<br>(0.039) | 0.008<br>(0.032) | 0.270***<br>(0.039) | 0.016<br>(0.028) |
| Population density in 1911 × log(LLD) | -0.155*<br>(0.085) | 0.432***<br>(0.044) | -0.010<br>(0.046) | 0.537***<br>(0.124) |
| Help to Buy (post-2015) x London dummy | 0.058***<br>(0.017) | 0.140***<br>(0.017) | 0.032**<br>(0.013) | 0.102***<br>(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).

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).[28] 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).

*Price-to-Rent Ratios (Full Sample)*

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 increased in an average LPA in response to an average local housing demand shock and the impact of the labor demand shock is stronger when regulatory (refusal rate) and physical (share developed land, altitude range) supply constraints are tighter, 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 are reported in columns (3) and (4) of Table 3. 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 km$^2$). [b)] PRP vs. market rent outliers (mean log market rent > 7.5, based on Figure 4) removed.

[28] As discussed below, we conduct several robustness checks that use a more refined approach. We use the 7.5 log points threshold in our baseline analysis because it is simpler. However, the results do not hinge on this choice.

*Heterogeneity over the Local Business Cycle*

According to our *Proposition 3 (iii)*, increased 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. As we document above (Panel B of Figure 5), the observed persistence in the labor demand shocks is much larger in the middle of a boom (or bust), whereas it is lower around turning points of the local business cycle (i.e., at the beginning of a local boom or bust and if a boom is overly long).

If participants in the housing market understand these relationships, the impact of a given labor demand shock on the price-to-rent ratio in our baseline regression should be driven mostly by periods where the degree of persistence is high – the 'boom' phases in Panel B of Figure 5. Conversely, if agents in the market were to become irrationally exuberant as a boom progresses, the relationship should grow stronger with each year the boom continues (i.e., during phases when booms become 'overly long'). To investigate these two competing propositions, we estimate separately the baseline specification for the sub-samples that capture the five phases of the local business cycle. Figure 6 displays the estimated coefficients for the supply constraint-interactions, along with 95% confidence intervals.

Although all interaction terms are significant at the beginning of the local business cycle, they become much larger as the local labor demand expands above its trend growth for at least two consecutive years (i.e., in the midst of a boom) – when the objective degree of demand shock persistence is highest. As the boom becomes 'overly long', the coefficients become smaller again, and the confidence bands widen – leading to mostly insignificant coefficients during this phase. This pattern lends strong support to our *Proposition 3 (iii)*, while it is not consistent with irrational exuberance.[29]

*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.[30]

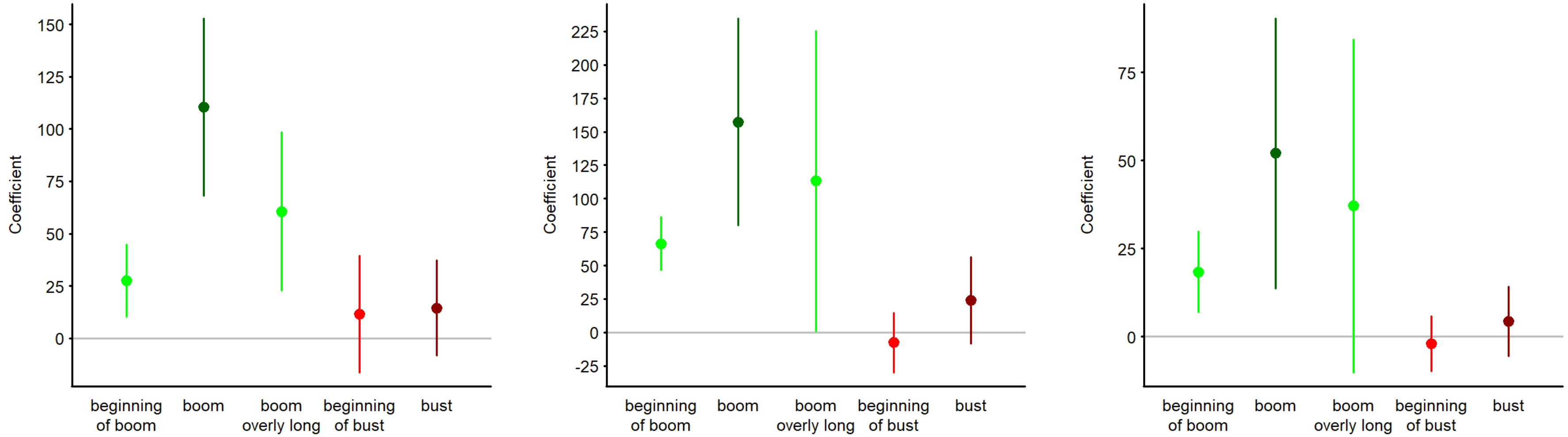
[29] We also ran regressions splitting the sample into LPAs with above- and below-median average persistence. The supply constraints interactions remain significant in both regressions, but they are considerably more important in the sub-sample of LPAs with above-median average persistence, consistent with *Proposition 3 (iii)*.

[30] 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

*Regression Coefficients of Interaction Effects for Different Phases of the Local Business Cycle*

*Panel A.* Log(LLD) × refusal rate | *Panel B.* Log(LLD) × share developed | *Panel C.* Log(LLD) × elevation range




*Notes:* The graphs display regression coefficients obtained from estimating the baseline specification (equation 7) separately by phase of the local business cycle. 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 6 (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*.[31] Table 6 reveals that the excluded instruments again correlate strongly and in expected ways with the endogenous supply constraint-measures.

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)] | Price-to-rent ratio 2SLS [a)] 1997-2018 [b)] |
| | $\triangle$LLD>0 | $\triangle$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 6. 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.

[31] We present corresponding results for house prices and rents in Table O-B1 of Online Appendix O-B. The results are qualitatively similar.

Table 6

*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.

### *3.5 Alternative Mechanisms*

While our findings are consistent with our proposed mechanism and inconsistent with irrational exuberance, several alternative explanations are also conceivable. We explore these one by one below and report 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 B1, 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.[32]

*Mortgage Financing Conditions*

Unobserved shocks to the relative (financing-)cost of homeownership could be correlated with changes in our measure of local housing demand, and lower costs or higher availability of mortgage credit could induce higher demand for owner-occupied housing relative to renting. To the extent 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.[33] 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 Appendix Table B2 we add the real rate of mortgage interest interacted with the supply constraint-measures (instrumented) as additional controls. In column (2), we repeat this exercise but use the spread measure interacted with the supply constraints (instrumented as well) instead.

Appendix Table B2 reveals that 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 (1.48) increases the difference in the price-to-rent ratio by only $1.48 \times 0.53 = 0.78$. In contrast, increasing the log labor demand by one within-standard deviation (0.05) has a much larger effect of $0.05 \times 51.7 = 2.59$. In a similar vein,

[32] 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.

[33] We use the Bank of England's quoted mortgage interest rate deflated by the Retail Price Index All Items Excluding Mortgage Interest (RPIX). Over our sample period, the real rate ranges between 3.91 and 8.59, while the spread ranges between 3.13 and 4.85, with standard deviations of 1.48 and 0.50, respectively.

decreasing the spread by one standard deviation (0.50 percentage points) increases the difference in the price-to-rent ratio by only $0.50 \times 0.60 = 0.30$, compared to 3.01 for a one-within-standard deviation increase of the log labor demand.

*Heterogeneous Impact of the Help to Buy Policy*

In Appendix Table B3, we test the robustness of our results to controlling more rigorously for the effects of the Help to Buy policy, which was introduced in England in 2013. As noted above, the policy provides a subsidy to homeownership. Although, in principle, the subsidy was not location-specific, except being more generous from 2016 onwards in the Greater London Authority, and the year fixed effects already control for its average impact on the price-to-rent ratio, differences in supply constraints could have led to differential impacts on house prices over space. We therefore define a second Help to Buy-dummy that is equal to one after 2012 and add the interactions of this dummy with the supply constraints measures to the regression.[34] The coefficients of the labor demand measure and its interactions remain qualitatively and quantitatively stable.

*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.[35] 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 B4 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 in England, rent stickiness in existing contracts is not an important phenomenon.

[34] The additional endogenous variables are instrumented by the interactions of the instrumental variables for the supply constraints with the Help to Buy dummy.

[35] 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.

*Can Global Investor Demand Explain the Increase in London's Price-to-Rent Ratio?*

Finally, we examine the hypothesis that global investor demand and other London-specific shocks may explain the relative increase of the price-to-rent ratio in London. To put an upper bound to the quantitative importance of these channels, we analyze the residuals of the baseline regression for the Greater London Authority. The residuals should capture the overall impact of all other relevant factors orthogonal to the local labor demand shocks, such as changing demand for real estate in London from global investors. Panel A of Figure C2 plots the prediction for London (black solid line) together with the year fixed effects (dashed grey line) and the average regression residual for LPAs located in Greater London (blue solid line). The dark red solid line represents the sum of the year fixed effects and the average residual, capturing the total impact of aggregate variables and London-specific shocks uncorrelated with the local labor demand shocks.

The blue solid line clearly shows that there is little room for global investor demand as an explanation for the substantial increase of London's price-to-rent ratio. The line hovers around zero since 2010. During the boom years before the Great Financial Crisis, the average residual in London was positive, but this reversed already before the peak in 2007. Overall, the net impact of other London-specific factors seems to be rather small.[36]

Panels B, C, and D display analogous graphs for the South East, 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 B5 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.

[36] 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. However, relative to the overall housing market of London, these markets are too small to influence much the development in the Greater London area.

*Choice of Rent Measure*

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

In Appendix Table B6, we then 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 (columns (1) to (3)).[37] 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. 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.

*Local Labor Demand Shock: A Placebo Test*

A third 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 correlated time series. To test this, we re-create the shift-share measure based on simulated

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

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.[38] 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 simulated 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.

A fourth 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 B7, we therefore replace the original labor demand-measure by 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 use the baseline regression from the preceding section (Table 4) to 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 ratio of a hypothetical location with relatively lax housing supply constraints.

## *4.1 Decomposition*

In Figure 7, we use the baseline regression to decompose the overall evolution of the price-to-rent ratio in selected locations – London, the South East, and the North East of England – into the impact of the aggregate component (the year fixed effects) and the effects of local labor demand shocks and their interactions with the housing supply constraints.[39] We select London because it experienced strong labor market shocks and has severely constrained housing supply,

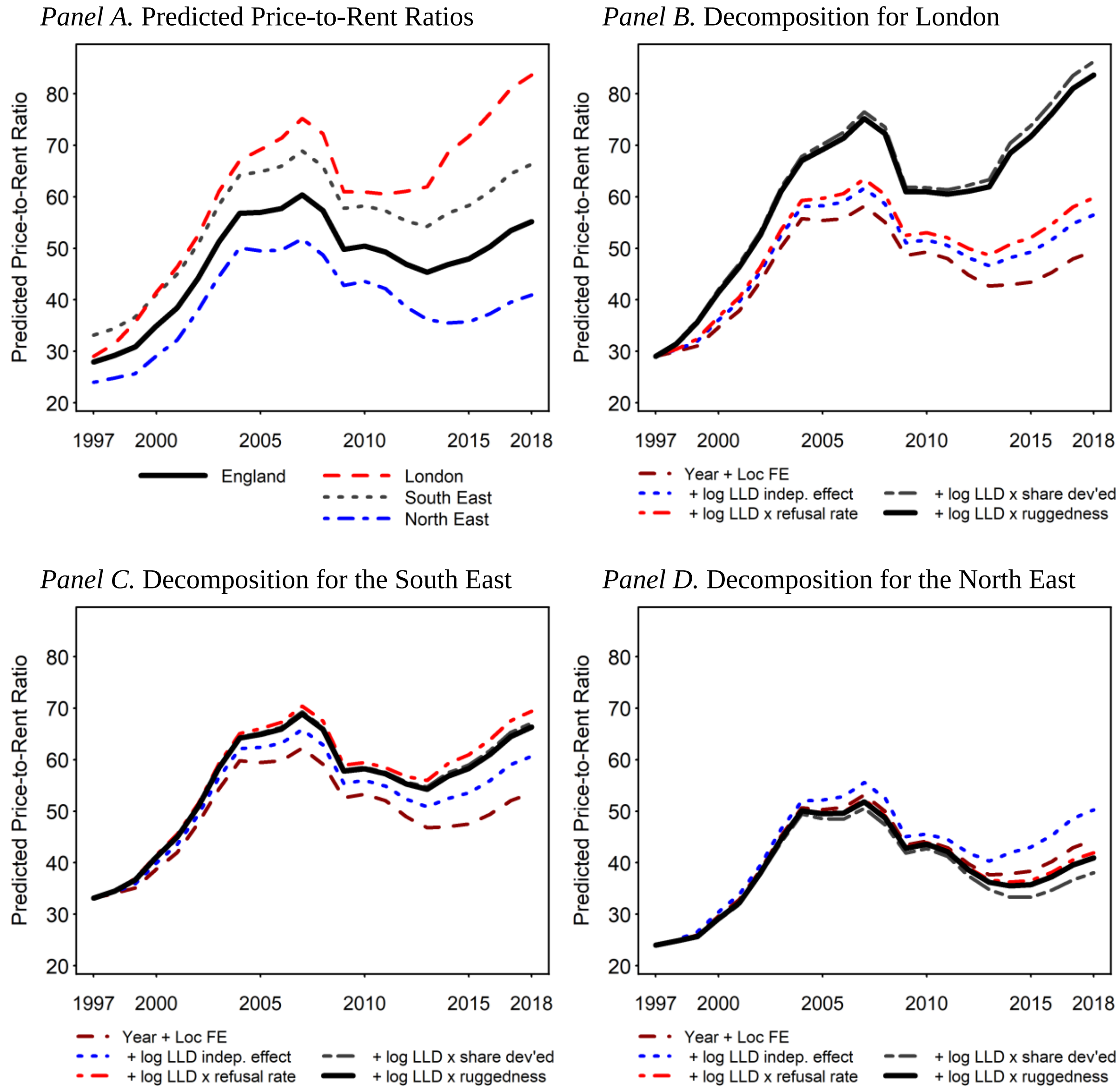
[38] The Akaike information criterion selects a lag order of 2 for the construction industry, and a lag order of 1 for all other industries.

[39] The predictions are based on the LPAs included in the baseline sample. Moreover, we weigh each LPA in the prediction by its share of households in the 2011 Census. The results are not sensitive to either of these choices. We report unweighted results for sections 4.1 and 4.2 in Figures O-C1 and O-C2 of Online Appendix O-C.

mainly due to a high share of developed land. The neighboring South East region is characterized by very tight regulatory constraints. Both regions are good examples of "location B" in Figure 3. The third region, the North East, has rather lax supply constraints, thus representing an example of "location A" in Figure 3.

Figure 7

*Predicted Price-to-Rent Ratios and Decomposition of the Price-to-Rent Ratio in London, the South East, and the North East of England*




*Notes:* All four graphs are based on the model displayed in Table 4. 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 shows the model-predicted price-to-rent ratios for England and for the Government Office Regions London, the South East, and the North East. Panel B decomposes the prediction for London (black solid line) into the impact of the fixed effects (dark-red dashed line), the independent effect of the local labor demand measure (blue dotted line), and its interaction effects with local regulatory restrictiveness (red dashed-dotted line) and the share developed land (grey long-dashed line). The difference between the grey long-dashed line and the black solid line represents the impact of the labor demand-ruggedness interaction term. Panels C and D display the respective graphs for the South East and the North East.

Panel A of Figure 7 displays the predicted price-to-rent ratios for England on average (black solid line) and for the three selected regions. The price-to-rent ratios differ markedly between regions and over time. The variation between locations is substantial, suggesting that the mechanism proposed in this paper is quantitatively important also relative to variation in price-to-rent ratios induced by macroeconomic variables. In Panel B, we decompose the prediction for London (solid black line) into the aggregate component (year fixed effects, dark red dashed line), the independent effect of local labor demand shocks (blue dotted line), the impact of the refusal rate interaction (red dashed-dotted line), and the impact of the share developed interaction (grey long-dashed line). The share developed interaction has the largest quantitative impact, clearly exceeding the aggregate component. The total effect of local labor demand shocks and their interactions with supply constraints represent 63% of the overall increase between 1997 and 2018, whereas the aggregate component explains the rest – 37% of the increase.

Panels C and D repeat this exercise for the South East and the North East of England. In the South East, the overall impact of the aggregate component is larger than the effect of local housing demand shocks in conjunction with local housing supply constraints, but the latter still account for a sizeable share of the overall increase (38%). Here, the main drivers are the refusal rate interaction and the independent impact of the labor demand shocks. In the North East, local labor demand shocks and their interactions with supply constraints are not important for explaining changes in the price-to-rent ratio. This fits nicely with the theoretical prediction for a location with comparably relaxed supply constraints. Our empirical model suggests that the local labor demand shocks and their interactions led to a slight decrease of the price-to-rent ratio between 1997 and 2018, thereby cushioning the overall increase due to macroeconomic and other common factors (as captured by the year fixed effects).

### *4.2 Counterfactual Exercise*

The independent effect of the local labor demand measure captures the impact of a local labor demand shock in an average location in England. This complicates the interpretation of the decomposition exercise: Arguably, the English planning system is one of the strictest planning systems – 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 Section 4.1 underestimates the importance of local housing supply constraints relative to countries with a higher average housing supply elasticity.

We therefore conduct an additional decomposition exercise that compares the three selected regions with a hypothetical 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). To rule out that differences in local labor demand shocks influence the results, we use the same labor demand shocks (i.e., the average labor demand shocks in England) for each location, including the hypothetical region, when calculating the predicted price-to-rent ratio. All differences between the hypothetical region and the selected location are then due to differences in housing supply constraints interacted with the common labor demand shocks.

Figure 8 shows the four hypothetical comparisons, separately for England (Panel A), London (Panel B), the South East (Panel C), and the North East of England (Panel D). We decompose the difference between the hypothetical place (dark-red dashed line) and the comparison region (black solid line) further into the impacts of the refusal rate interaction (red dashed-dotted line) and the share developed interaction (grey long-dashed line). The difference between the grey long-dashed and the black solid line represents the impact of the elevation range interaction.

The four graphs in Figure 8 suggest 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. Relative to this place, the impact of local labor demand shocks in conjunction with local supply constraints on the price-to-rent ratio is very substantial for London and the South East, but also for an LPA with average levels of supply constraints. This suggests that the fixed effects capture to a substantial degree the impact of aggregate housing demand in conjunction with aggregate supply constraints, besides interest rates and other credit conditions.

# 5 Conclusions

The importance of housing market dynamics for macroeconomics has become clear in the wake of the Great Financial Crisis. While this financial crisis can be associated with one boom-bust cycle, another housing-related crisis has been slowly brewing over decades: Rising house prices and rents have triggered a crisis of housing affordability. This crisis, while also global in nature, is particularly acute in large desirable and tightly supply constrained – so called 'superstar' – cities. This is perhaps why it has particularly raised the interest of urban economists.

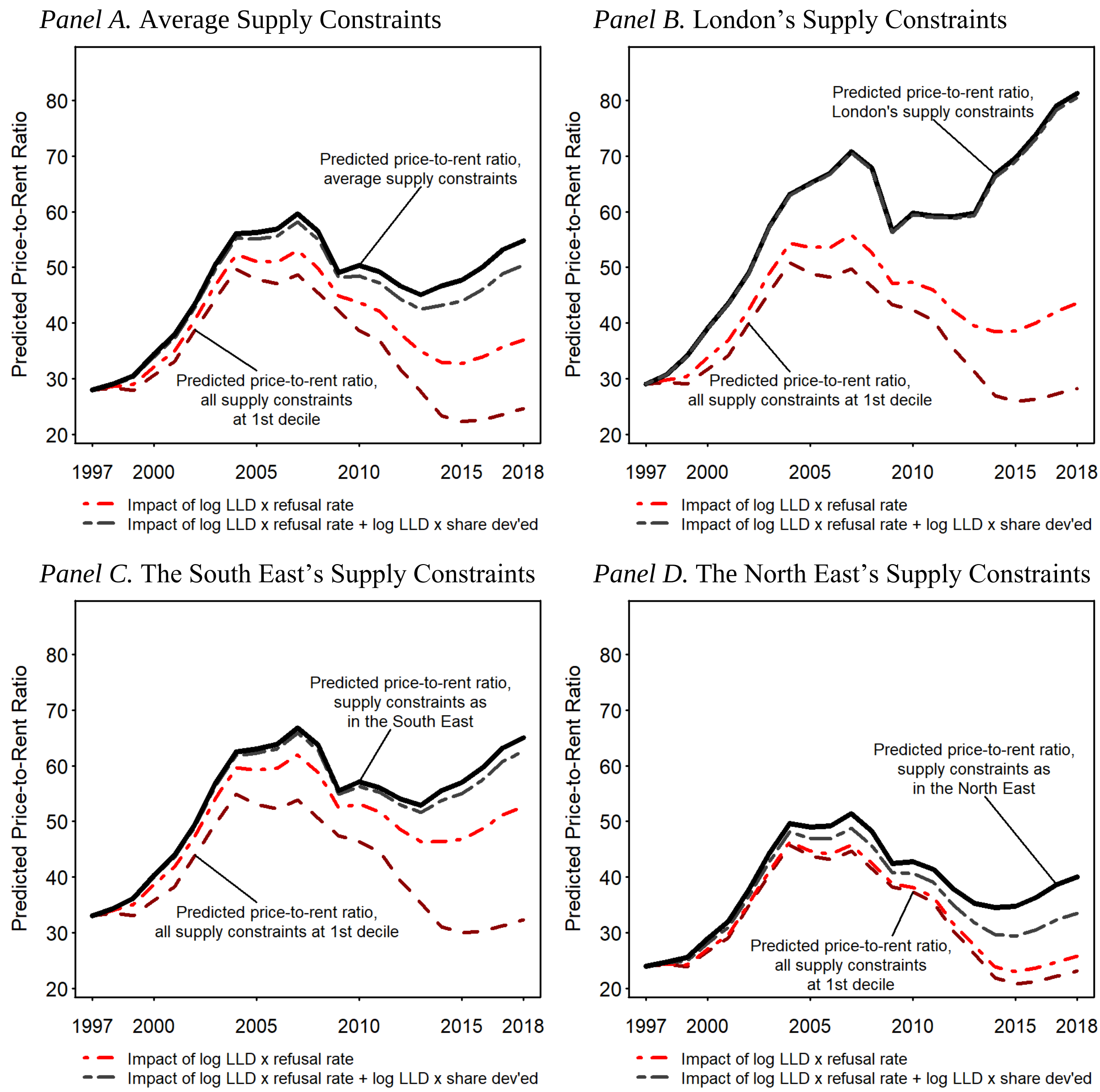
The mounting housing affordability crisis is (again) hitting younger and lower income households the hardest, especially those who are aspiring to become homeowners. It is arguably contributing to political unrest, with the 2019-20 Hong Kong protests being just the latest and most prominent example.

The underlying causes of the 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 in particular is highly policy relevant: To what extent are the rising house prices driven by land use planning-induced housing supply shortages?

One of the most striking stylized facts in this context is that in many places, particularly in superstar cities, over the last two decades, house prices have risen much more strongly than rents. This fact has been invoked by many as evidence that the 'housing supply shortage hypothesis' is misplaced, with the argument being that supply constraints should drive up prices and rents equally. Falling real interest rates, changing credit conditions, unrealistic expectations, rising inequality in conjunction with segmented housing markets, or global investor demand for superstar cities are offered by these proponents as alternative explanations.

Figure 8

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




*Notes:* All four graphs are based on the model displayed in Table 4, employing the average labor demand shocks in England during the sample period. LPA-level predictions were aggregated to Government Office Regions, employing the number of households in each LPA in 2011 (Census) as weights. Panels A-D show model-predicted price-to-rent ratios for (household-weighted) average supply constraints in England, London, the South East, and the North East (black solid line). All four graphs also show the predicted price-to-rent ratio for a location where all supply constraints are set to the 10% sample quantile (dark-red dashed line), shifted vertically to match the 1997 price-to-rent ratio of the respective location. The red dashed-dotted lines represent the impact of changing the refusal rate from the 10% quantile to the counterfactual location's refusal rate. The dashed grey line adds the impact of changing the share developed from the 10% quantile to the respective location's share developed. The remaining difference to the black solid line represents the impact of changing the elevation range from the 10% quantile to the respective location's elevation range.

In this study we provide a simple theory – tight supply constraints in conjunction with serially correlated demand shocks – that does not rely on any of these alternative explanations. It can not only to a good extent explain the rising price-to-rent ratio over the last two decades, especially in superstar cities, but also several other regularities in housing market dynamics across the globe. These include the observations that (i) the increase in the price-to-rent ratio tends to be most pronounced in the most desirable and supply constrained 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 thriving places and superstar cities* like London 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, are crucial drivers explaining the price-to-rent-ratio dynamics. The year fixed effects are a 'black box' that are likely to capture real interest rates, but also plausibly the country's credit conditions, and 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, $\varepsilon$:

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

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. $\beta$,

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

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

Appendix Table B1

*Controlling for Local Income Inequality*

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | Price-PRP rent ratio, P80/P20 | Price-PRP rent ratio, P80/P20 | Price-PRP rent ratio, approx. Gini | Price-PRP rent ratio, approx. Gini |
| Log(local labor demand) | 39.520*** | 37.480*** | 36.158*** | 36.106*** |
| | (10.995) | (11.353) | (11.501) | (11.250) |
| Av. refusal rate of major residential projects × log(local labor demand) | 61.853*** | 62.339*** | 62.103*** | 61.362*** |
| | (9.268) | (9.189) | (9.580) | (9.034) |
| Share of developable land developed in 1990 × log(local labor demand) | 80.630*** | 81.969*** | 81.789*** | 80.620*** |
| | (18.024) | (17.838) | (18.247) | (16.819) |
| Range between highest and lowest altitude × log(local labor demand) | 22.435*** | 22.731*** | 22.119** | 21.711*** |
| | (8.486) | (8.478) | (8.607) | (8.116) |
| Help to Buy (post-2015) x London dummy | -1.101 | -1.470 | -1.263 | -1.153 |
| | (3.195) | (3.265) | (3.213) | (3.007) |
| Local income inequality | -1.744 | -2.437** | 1.490 | 2.766 |
| | (1.094) | (1.100) | (3.573) | (3.579) |
| Av. refusal rate of major residential projects × local income inequality | | 0.527 | | 1.869 |
| | | (1.354) | | (5.322) |
| Share of developable land developed in 1990 × local income inequality | | -3.941** | | 18.229*** |
| | | (1.757) | | (6.829) |
| Range between highest and lowest altitude × local income inequality | | 1.091 | | -1.087 |
| | | (1.404) | | (5.198) |
| LPA FEs | Yes | Yes | Yes | Yes |
| Year FEs | Yes | Yes | Yes | Yes |
| Observations | 6,830 | 6,830 | 6,735 | 6,735 |
| Number of LPAs | 344 | 344 | 344 | 344 |
| Kleibergen-Paap F | 11.07 | 5.30 | 10.71 | 6.31 |

*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 B2

*Mortgage Financing Conditions*

| | (1) | (2) |
|---|---|---|
| | Real Mortgage Interest Rate | Mortgage Rate Spread |
| Log(local labor demand) | 30.192** | 30.032** |
| | (12.581) | (12.514) |
| Av. refusal rate of major residential projects × log(local labor demand) | 51.646*** | 61.115*** |
| | (7.048) | (8.376) |
| Share of developable land developed in 1990 × log(local labor demand) | 68.621*** | 85.019*** |
| | (14.251) | (16.569) |
| Range between highest and lowest altitude × log(local labor demand) | 18.337*** | 21.447*** |
| | (7.033) | (8.056) |
| Help to Buy (post-2015) x London dummy | -0.069 | -2.080 |
| | (2.903) | (2.916) |
| Av. refusal rate of major residential projects × Real Mortgage Interest Rate | -0.525*** | |
| | (0.141) | |
| Share of developable land developed in 1990 × Real Mortgage Interest Rate | -0.506*** | |
| | (0.192) | |
| Range between highest and lowest altitude × Real Mortgage Interest Rate | -0.226** | |
| | (0.090) | |
| Av. refusal rate of major residential projects × Mortgage Interest Rate Spread | | -0.600 |
| | | (0.401) |
| Share of developable land developed in 1990 × Mortgage Interest Rate Spread | | 0.911* |
| | | (0.474) |
| Range between highest and lowest altitude × Mortgage Interest Rate Spread | | -0.839*** |
| | | (0.289) |
| LPA FEs | Yes | Yes |
| Year FEs | Yes | Yes |
| Observations | 7,555 | 7,555 |
| Number of LPAs | 344 | 344 |
| Kleibergen-Paap F | 5.18 | 5.23 |

*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.

Appendix Table B3

*Help to Buy*

| | (1) |
|---|---|
| | Help to Buy Interactions |
| Log(local labor demand) | 40.067*** |
| | (9.942) |
| Av. refusal rate of major residential projects × log(local labor demand) | 51.856*** |
| | (7.930) |
| Share of developable land developed in 1990 × log(local labor demand) | 47.401*** |
| | (11.840) |
| Range between highest and lowest altitude × log(local labor demand) | 31.935*** |
| | (7.597) |
| Help to Buy (post-2015) x London dummy | -1.498 |
| | (2.884) |
| Av. refusal rate of major residential projects × Help to Buy (post-2012) | 1.010 |
| | (0.635) |
| Share of developable land developed in 1990 × Help to Buy (post-2012) | 4.344*** |
| | (0.900) |
| Range between highest and lowest altitude × Help to Buy (post-2012) | -1.148** |
| | (0.485) |
| LPA FEs | Yes |
| Year FEs | Yes |
| Observations | 7,555 |
| Number of LPAs | 344 |
| Kleibergen-Paap F | 7.17 |

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 post-2012 dummy are instrumented by the interactions of the post-2012 dummy with the instruments discussed in Section 3.2.

Appendix Table B4

*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 B5

*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 projects × log(local labor demand) | 65.038*** | 62.086*** | 48.848*** | 52.034*** | 39.026* | 69.331*** |
| | (13.285) | (8.927) | (10.704) | (11.122) | (22.928) | (14.161) |
| Share of developable land developed in 1990 × log(local labor demand) | 84.418*** | 81.511*** | 79.825*** | 80.872*** | 74.630*** | 89.172*** |
| | (22.250) | (18.045) | (17.349) | (17.170) | (21.323) | (23.060) |
| Range between highest and lowest altitude × log(local labor demand) | 25.101** | 23.763*** | 22.262*** | 22.878*** | 19.585* | 27.256** |
| | (10.324) | (8.635) | (8.032) | (7.995) | (10.039) | (10.750) |
| Help to Buy (post-2015) × London dummy | -1.777 | -1.192 | -0.699 | -0.938 | 0.408 | -2.727 |
| | (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 B6

*Robustness Checks for Selection of Rent Measure*

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Dependent variable | | Price-<br>PRP rent ratio | | Price-<br>market rent<br>ratio | Price-<br>PRP rent<br>ratio |
| LPA-level correlation of<br>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 B7

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

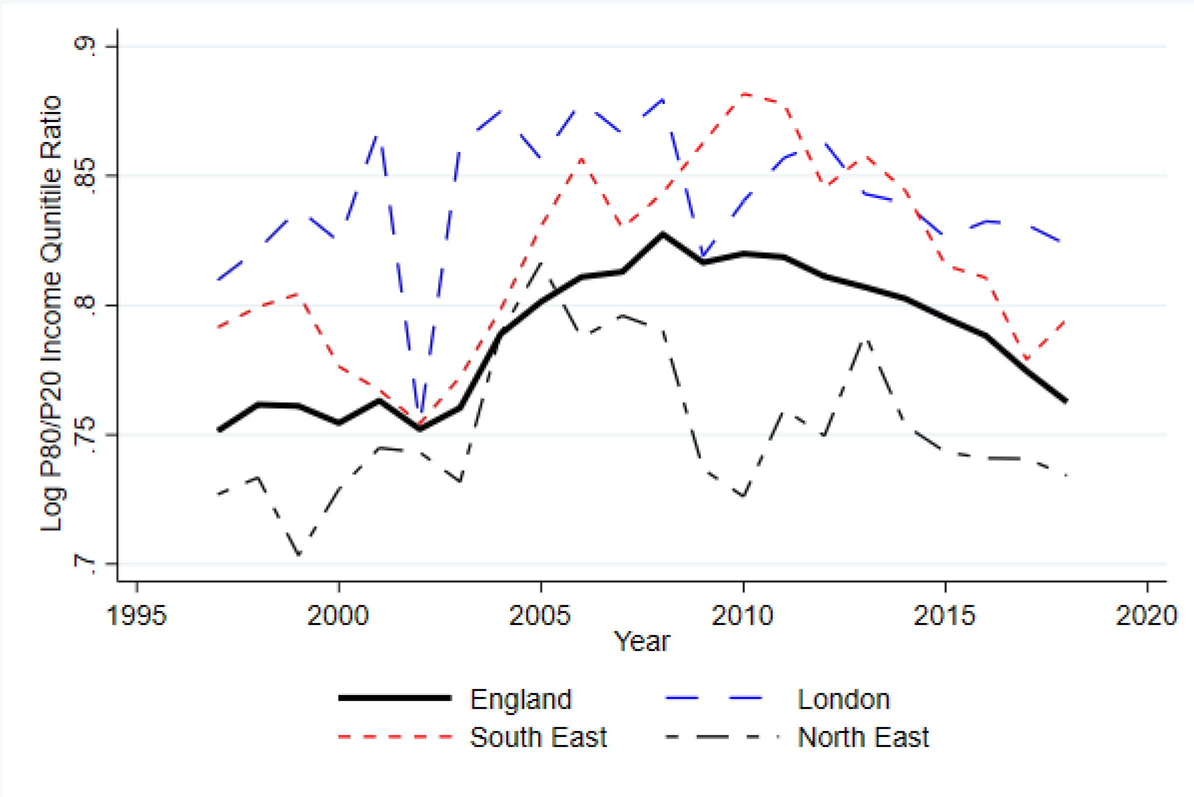
| | (1) |
|---|---|
| | LLD w/o banking and real estate services |
| Log(adjusted local labor demand) | 39.374*** |
| | (10.617) |
| Av. refusal rate of major residential projects × log(adjusted local labor demand) | 59.973*** |
| | (8.786) |
| Share of developable land developed in 1990 × log(adjusted 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

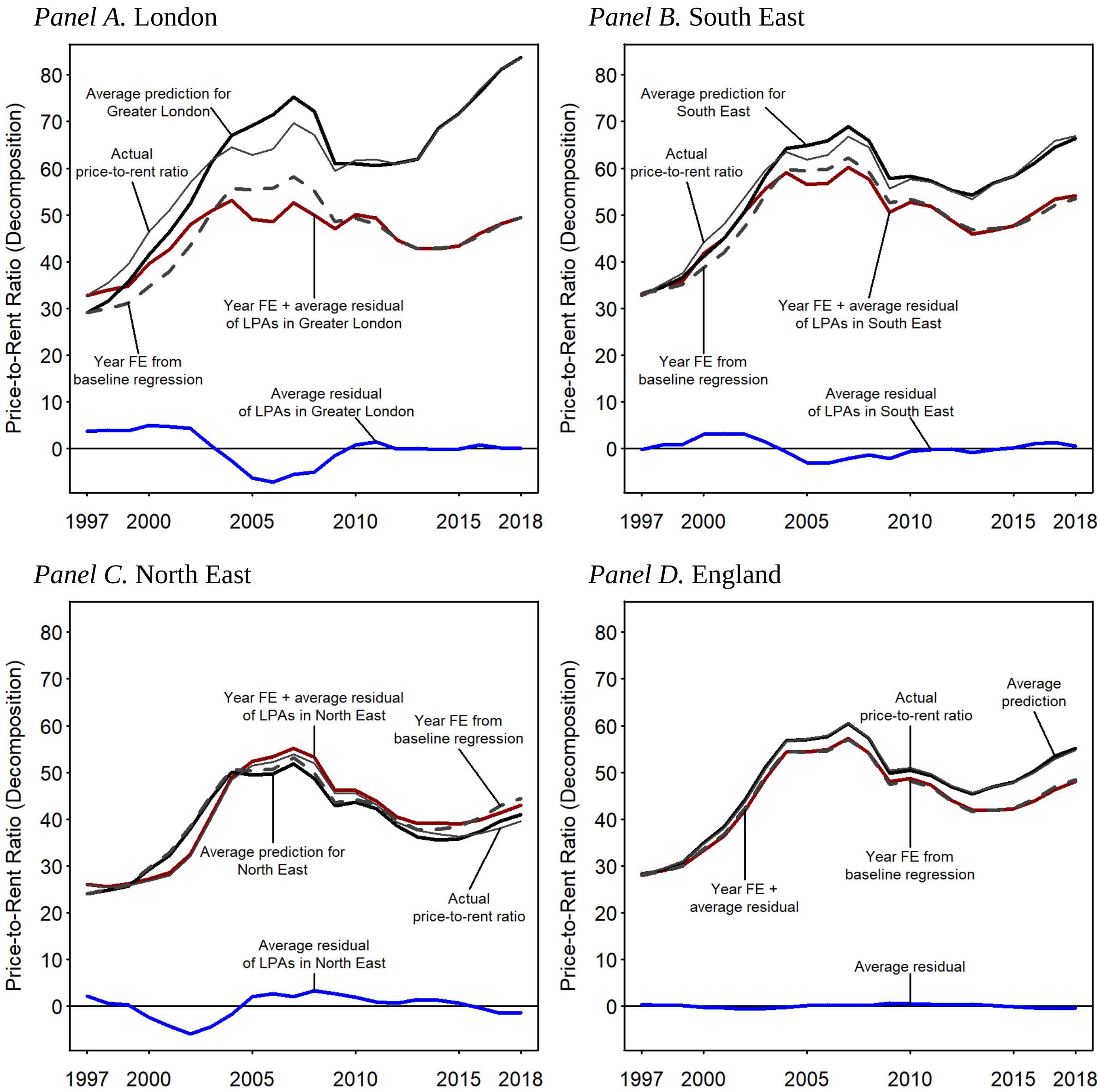
*Income Inequality in England, London, the South East, and the North East of England, 1997-2018*




*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

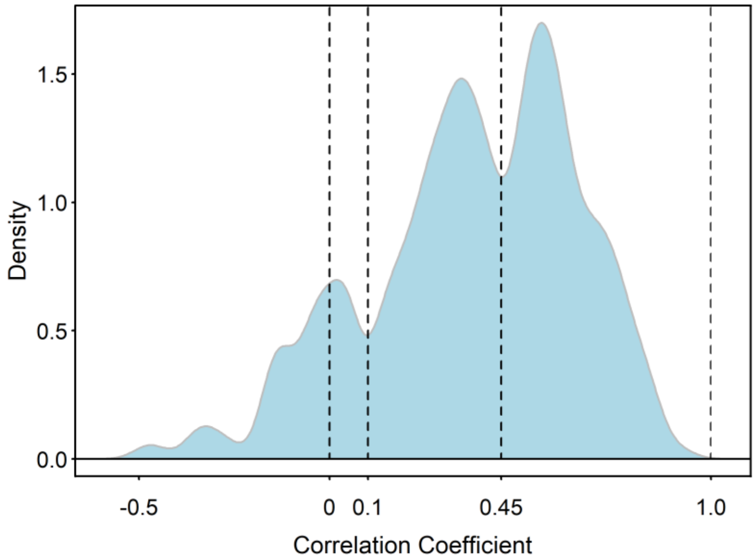
*Residual Variation in the Price-to-Rent Ratio by Region and in England*




*Notes*: The graphs decompose for each region the average predicted price-to-rent ratio (black solid line) into the year fixed effects (dashed grey line) and the impact of the local labor demand measure and its interactions with the local housing supply constraints (difference between the solid black line and the dashed grey line). Moreover, the thin grey line represents the actual average price-to-rent ratio and the blue line represents the average regression residual. The dark red solid line represents the sum of the year fixed effects and the average residual, capturing the total impact of aggregate variables and region-specific shocks, such as credit conditions and demand for London real estate from global investors. Panels A to C refer to the government office regions of London, the South East, and the North East. Panel D refers to England as a whole.

Appendix Figure C3

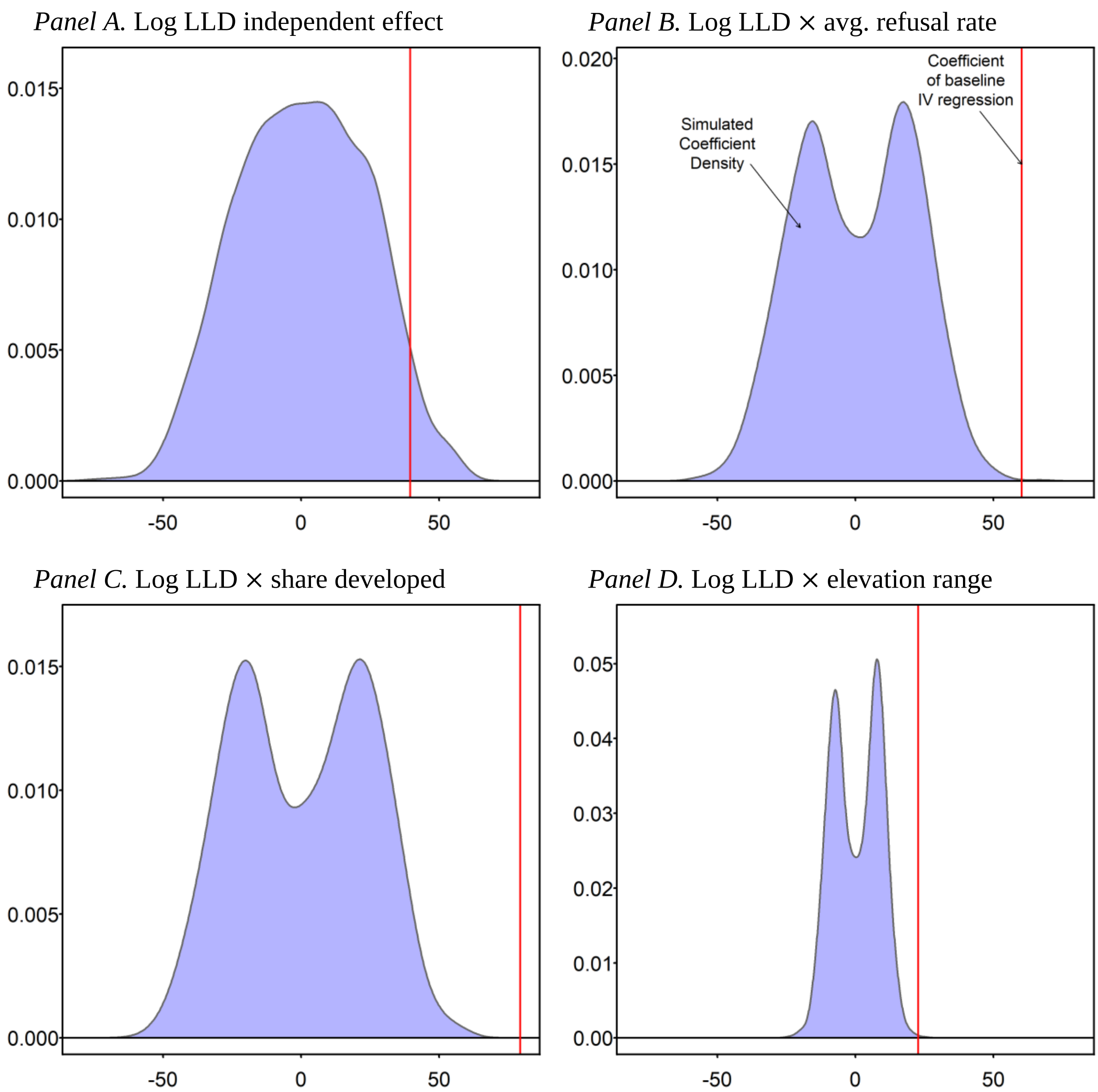
*Distribution of Correlation between Changes in Market Rents and PRP Rents at LPA-Level*




*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*




*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 1974-1994 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, 1995-2018, to the adjusted 1974-1994 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<br>England (Census) | <br>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).

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, 2010-2018.* 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-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, 1997-2018.* The uk.gov Table 704 of the UK Housing Statistics reports mean rents charged by Private Registered Providers (PRP), by year (1997-2018), 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: Ministry of Housing, Communities and Local Government), 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)] | Prices 2SLS [a)] 1997-2018 [b), c)] | Rents 2SLS [a)] 1997-2018 [c)] | Rents 2SLS [a)] 1997-2018 [c)] |
| | △LLD>0 | △LLD≤0 | △LLD>0 | △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 |

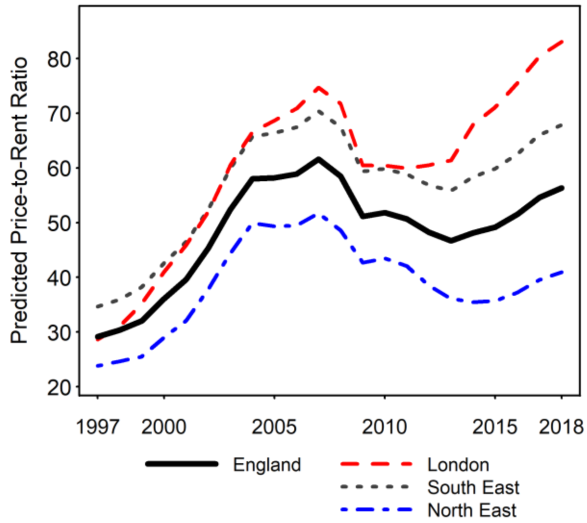
*Notes:* Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. [a)] First stage results are reported in Table 6 in the main text. Instruments include: Share of greenbelt land in 1973, change in delay rate b/w 1994–96 & 2004–06, share of votes for Labour in 1983 General Election, 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: Decomposition and Counterfactual Graphs— Without Weighting by the Number of Households

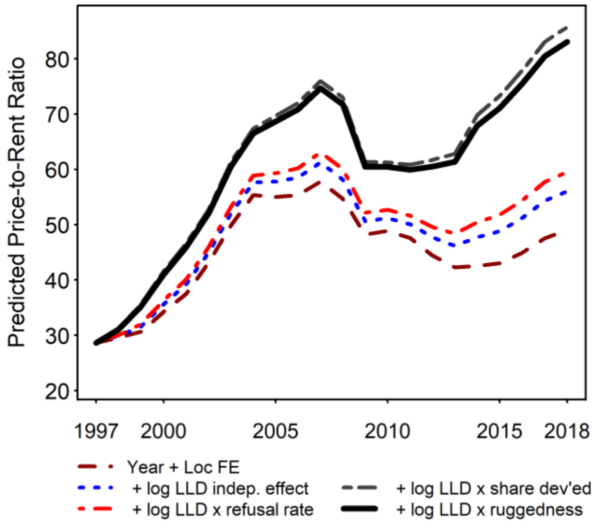
Figure O-C1

*Predicted Price-to-Rent Ratios and Decomposition of the Price-to-Rent Ratio in London, the South East, and the North East of England, Unweighted*

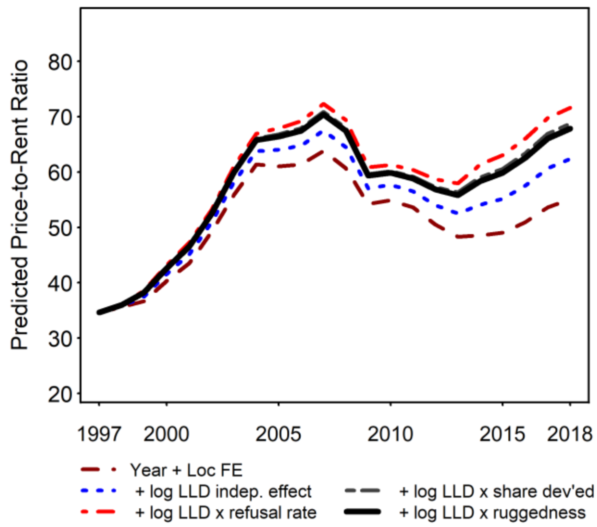
*Panel A.* Predicted Price-to-Rent Ratios



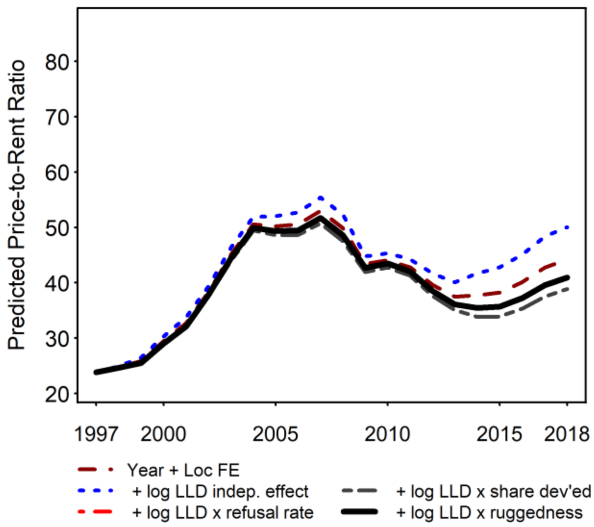

*Panel B.* Decomposition for London




*Panel C.* Decomposition for the South East



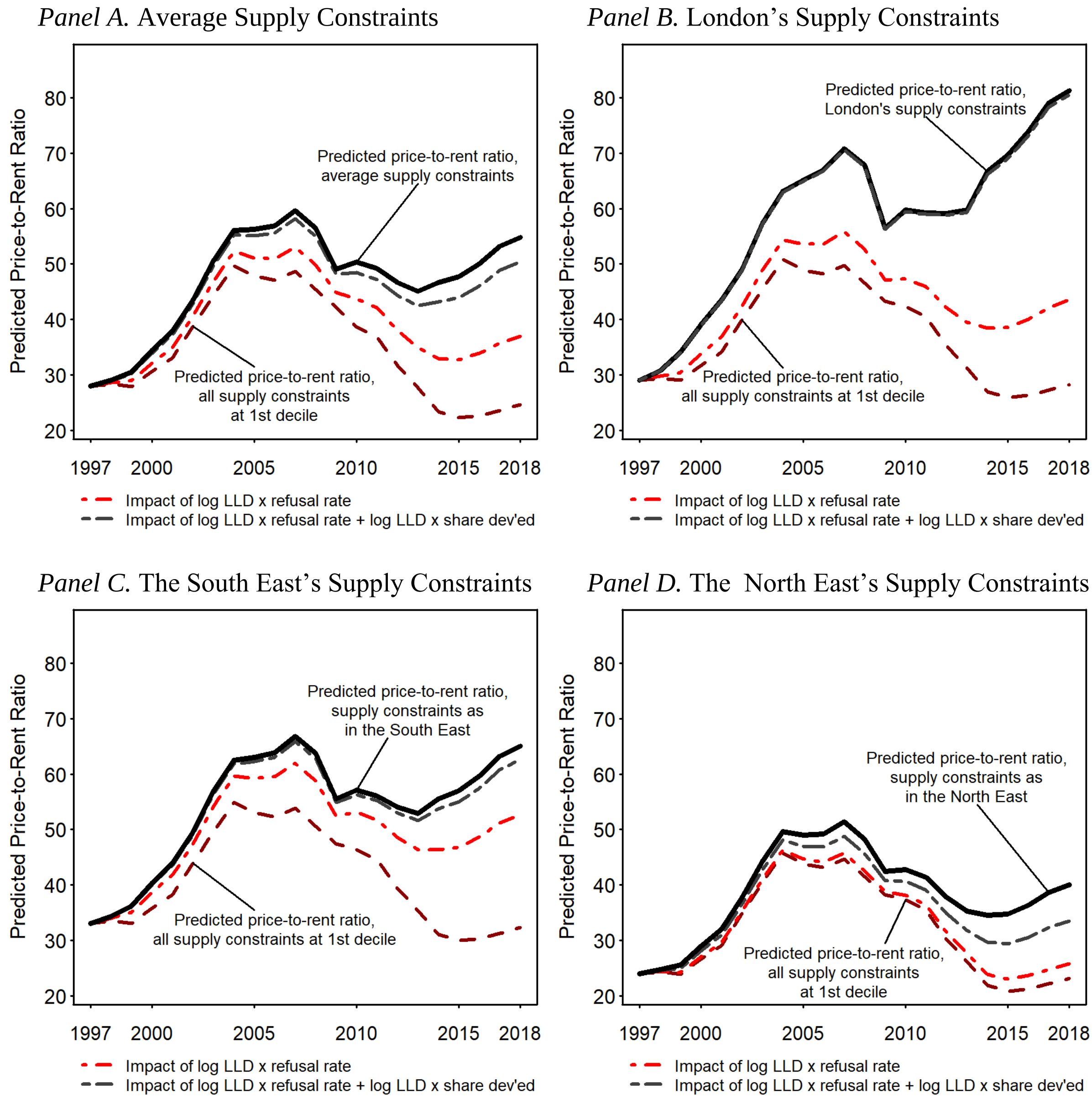

*Panel D.* Decomposition for the North East




*Notes:* All four graphs are based on the model displayed in Table 4. The model was used to compute LPA-level predictions that were aggregated to Government Office Regions. Panel A shows the model-predicted price-to-rent ratios for England and for London, the South East, and the North East. Panels B-D decomposes the predictions for London, the South East, and the North East (black solid line) into the fixed effects (dashed dark red line), the independent effect of the local labor demand measure (blue dotted line), and its interaction effects with the regulatory restrictiveness (red dashed-dotted line), and with the share developed (grey long-dashed line). The difference between the grey long-dashed line and the black solid line represents the impact of the labor demand-ruggedness interaction term.

Figure O-C2

*Counterfactual Decomposition Relative to a Location with Supply Constraints at the 10% Sample Quantiles, Unweighted*




*Notes:* All four graphs are based on the model displayed in Table 4, employing the average local labor demand shocks in England during the sample period. The predictions were aggregated to Government Office Regions. Panels A-D show model-predicted price-to-rent ratios for (household-weighted) average supply constraints in England, London, the South East, and the North East (black solid lines). All graphs also show the predicted price-to-rent ratio for a location where all supply constraints are set to the 10% sample quantile (dark red dashed lines), shifted vertically to match the 1997 price-to-rent ratio of the location. The red dashed-dotted lines represent the impact of changing the refusal from the 10% quantile to the counterfactual location's refusal rate. The dashed grey line adds the impact of changing the share developed from the 10% quantile to the respective location's share developed. The remaining difference to the black solid line represents the impact of changing the elevation range from the 10% quantile to the respective location's elevation range.

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