Regulating housing vacancies away?

The paradoxical effects of mismatch*

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Abstract

Policy makers use the existence of empty houses as an argument for being more restrictive in their planning policies. So higher vacancy rates tend to trigger tighter restrictions on the supply of land. Such tighter restrictions lead to higher prices and, because of the incentives this creates for occupying housing, to lower housing vacancies ('opportunity cost effect'). There is, however, a second effect ignored by planners: more restrictive planning policies impede the matching process in housing markets so leading to higher vacancies ('mismatch effect'). Which of these two forces dominates is an empirical question. This is our focus here. Addressing potential reverse causation and other endogeneity concerns, we use a unique panel data set on land use regulation for 350 Local Authorities in England from 1981 to 2011. Our results show that tighter local planning constraints increase local housing vacancy rates, suggesting that the mismatch effect dominates. A one standard deviation increase in local regulatory restrictiveness causes the average local vacancy rate to increase by about 0.9 percentage points (23 percent). The same increase in local restrictiveness also causes an 8.5 percent rise in average commuting distances. The results are economically meaningful and show that pointing to the existence of vacant houses as a reason for being more restrictive in allocating land for housing is counterproductive.

JEL classification: R13, R38

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1. Introduction

To an economist it might seem self-evident that vacancies in the housing stock are a natural feature of how any market must work. There even are 'uneaten' apples in a well-functioning fruit market. The labour market is very much more comparable to the housing market and virtually all mainstream economists expect to observe at least frictional unemployment when the labour market is in equilibrium (see Pissarides, 1985; Mortensen and Pissarides, 1994; Pissarides, 1994). It is the same in any normally functioning housing market. In equilibrium there must be vacant houses as people move and 'house-hunt', as people die or houses wait to be demolished and sellers wait to find a buyer (Han and Strange, 2015).

But this view is often not shared by those who design buildings and influence urban policy or with those who plan housing supply – at least in England. Even in what was then one of the least restrictive English Regions, the East Midlands, in calculating how much land should be allocated for housing to meet their estimate of their region's 'housing needs', planners argued that they could allocate less land because they assumed they would reduce the number of vacant homes:

> '*The annual average housing provision reflects a number of factors, transactional vacancies in new stock (about 2%) add 7,000 to the requirement, but offset against that is an assumption that vacancies in the existing stock should be reduced by a half per-cent, which will bring 8,600 dwellings back into use.*' (Government Office for the East Midlands, 2005, Appendix 4, p. 91).

A logically equivalent assertion was made by the well-known architect and advocate of urban density, Lord Rogers, in arguing against the desirability of allowing offices to be converted to housing to help with London's housing supply: '…why should we rush to convert office blocks when we already have three-quarters of a million homes in England lying empty…' (Lord Rogers of Riverside, 2013).

It is surely true that using ones stock of capital more intensively is a way of increasing efficiency. That is just how the cut price airlines operate: they keep their seats full and their aircraft in the air. They, however, had an analysis of how to reduce the time aircraft spent on the ground and policies to achieve it. They did not just assume planes would spend more of their lives in the air. Unless we understand why houses are vacant we cannot hope to reduce the number of vacant houses just by being more restrictive. To help improve our understanding of the factors which determine vacancy rates in the housing market, this paper investigates the causal, albeit reduced-form, impact of regulatory restrictiveness, controlling for other determinants of vacancy rates.

This is important because restrictions on housing supply, imposed by planning, may have two opposing effects.[1] The first of these we call the 'opportunity cost effect'. Tighter restrictions on supply imply fewer available houses and therefore more demand pressure for existing

[1] Regulation may have more than two effects. We discuss one potential additional mechanism – a real options argument – in Section 3.5. If greater restrictiveness led to greater price volatility then under certain assumptions this might induce owners to postpone renting or selling their properties, implying a higher vacancy rate (Grenadier, 1995; 1996). Empirically, however, we can find no evidence that such a mechanism plays a significant part in explaining what we find.

homes, increasing house prices and thus the opportunity cost of keeping housing empty. This will lead to a lower vacancy rate all else equal. If this ‘opportunity cost effect’ was the only effect at work, tighter supply constraints should unambiguously lower vacancy rates.

There is however a second effect, which we refer to as the ‘mismatch effect’. Tighter supply constraints not only reduce supply of new houses but also influence the composition and adaptability of the bundle of attributes of both the housing stock and new build. Over time the structure of households’ demand for housing attributes changes because incomes rise, the demographic structure of the population changes and preferences themselves may change. For example, as real incomes rise so does the demand for certain attributes depending on the varying income elasticity of demand for them.[2] In addition there may be demographic changes such as an increase in the proportion of single adults, which mean that market preferences change.

If the attributes of the housing stock, as a consequence of planning constraints, cannot, or can only more slowly adjust to these changes on the demand side, matching the demand for housing attributes with the supply of those available will inevitably become more difficult. Hence, in line with Wheaton (1990), mismatched households may have to reside longer in a less restrictive housing market, while searching in a more restrictive market, implying a relatively lower vacancy rate in the less restrictive market and a higher vacancy rate in the more restrictive one. Mismatched households may also have to search further afield for a suitable home; they become mismatched on the locational characteristics of houses implying longer commutes.

Our aim in this paper is to determine the net effect of these two opposing forces – the opportunity cost effect vs. the mismatch effect – in order to identify the role that regulatory restrictiveness plays in determining the vacancy rate in local housing markets. To do so, we analyse panel data on housing vacancies for 350 Local Authorities (LA) in England from 1981 to 2011. One key concern in this analysis is the endogeneity of local planning restrictiveness. The stylised fact that policy makers and local planners may respond to higher vacancy rates by restricting supply suggests possible reverse causation. Regulatory constraints may also be endogenous to unobserved demand factors (Hilber and Robert-Nicoud, 2013; Davidoff, 2014) and those demand factors may directly affect vacancy rates. To account for possible reverse causation and omitted variable bias and thus identify the causal effect of regulatory restrictiveness, we employ an instrumental variables strategy by exploiting a specific feature of the British voting system; the substantial ‘randomness’ of seats won (or lost) beyond the vote share – the outcome of the ‘first past the post’ system. That is, we use the share of Labour seats in LAs (which implement planning policies in England), controlling for the share of Labour votes in a flexible way, as an instrumental variable to identify local planning refusal rates.

Our empirical findings are as follows. First, when we naively look at cross-sectional data, we find a negative relationship between more restrictive local planning and local vacancy rates. That is, the relationship superficially appears to confirm the planners’ assumptions. However,

[2] The income elasticity of demand for space both inside houses and in gardens seems to be particularly strong: Cheshire and Sheppard (1998) estimate an elasticity of close to two.

if we (*i*) use first differencing and so control for time-invariant unobservable characteristics, (*ii*) properly account for the endogeneity of restrictiveness by instrumenting for it and (*iii*) control for other relevant factors, more restrictive places have a significantly – and substantially – *higher* housing vacancy rate. That is the underlying causal relationship appears to be exactly the opposite to that which planners assume. Based on our most rigorous empirical specification, a one standard deviation increase in local regulatory restrictiveness causes the average local vacancy rate to increase by about 0.9 percentage points (23 percent).

We further find that tighter local restrictions create a particular mismatch – a 'spatial mismatch'. Workers with jobs in LAs which are more restrictive have to search for housing they can afford and match their preferences further afield; so they are more likely to be locationally mismatched and have to commute further. Using a similar approach to that used for investigating the underlying relationship between the vacancy rate and restrictiveness we find that a one standard deviation increase in local regulatory restrictiveness causes an increase in average commuting distance of some 8.5 percent.

Our findings, therefore, strongly suggest that tighter local planning restrictiveness not only leads to less efficient housing market matching but this effect dominates the opportunity cost effect, resulting in higher local vacancy rates overall and longer average commutes. Hence, local efforts to reduce the number of vacant homes by imposing supply restrictions have the paradoxical effect of increasing the local vacancy rate while at the same time making housing less affordable and causing a net welfare loss via additional commuting. We proceed as follows. In the next section we discuss the relationship between land use regulation, housing market search and matching and the resulting vacancy rate in more depth. We then describe our data and set out our main results. The final section draws conclusions.

2. Land use regulation, housing market search and vacant housing

The price of housing services is a function of both demand and supply in the relevant local markets. Various empirical studies document a positive effect of regulatory restrictiveness on house prices (Cheshire and Sheppard, 2002; Glaeser and Gyourko, 2003; Glaeser et al., 2005a; b; Quigley and Raphael, 2005; Ihlanfeldt, 2007; Hilber and Vermeulen, forthcoming).

What these studies do not consider is the fact that, on the seller's side, it takes time to sell a house and, on the buyer's side, search for a new house is costly too. These search frictions lead to housing vacancies (Merlo and Ortalo-Magné, 2004; Han and Strange, 2015). It has been documented – and our data also suggests – that housing vacancies are not constant across space and time and depend on the characteristics and preferences of households living in a housing market, as well as on characteristics of the location (Rosen and Smith, 1983; Gabriel and Nothaft, 2001). However, the impact of land use restrictions on housing vacancies has not yet been studied.

In the context of this paper we use data on local jurisdictions – in Britain, Local (Planning) Authorities (LAs) – which we refer to as local housing markets.[3] Households often search in a local housing market while still living in another local market, for example due to changes in where they work (Mulalich et al., forthcoming; Koster and Van Ommeren, 2015). Within any LA, the supply of housing units is the outcome of construction companies' building costs (including the cost of land) relative to expected future prices; the existing stock of houses and the degree of regulatory restrictiveness exercised by the LA. The supply of characteristics of housing is the result of both the characteristics of new build housing and the adaptation of the characteristics of the existing stock.

The regulatory restrictiveness influences the rate and characteristics of new construction and the flexibility of the existing stock since both new construction and significant changes to the characteristics of existing houses typically require 'development control' permission. That is, all proposals which fall within the legal definition of 'development', such as any change of legally defined 'use category', a qualifying addition of living space or a garage or building a new house or flat, have to be submitted to the LA for decision. This is the responsibility of the LA's Planning Committee made up of locally elected politicians. This decision making process tends to be politicised and unlike a Zoning or Master Planning system, such as in force in the US or in most of Continental Europe, decisions are not very predictable.

The degree of restrictiveness also influences the composition and adaptability of the bundle of attributes of existing housing available to meet the changing preferences of those seeking houses. This is because the more restrictive an LA is, the more difficult it will be to change the characteristics of a given house and the less flexibility will there be on the characteristics – including location – of new build houses: for example someone searching for a house with a good local school may require more bedroom space but it will be more difficult to obtain permission to convert attics into living space in more restrictive LAs. Equally, developers may have greater difficulty expanding the supply of family housing near better schools. In other words, all else equal, in more restrictive markets both the supply of new houses and the characteristics of the existing stock will be less well adapted to the structure of demand for housing characteristics.

As noted in the introduction, planning induced housing supply restrictions will have two opposing effects on the rate of the vacant housing stock: an 'opportunity cost effect' and a 'mismatch effect'. The opportunity cost effect works via restrictions of supply reducing the availability of land for development (see for example Cheshire and Sheppard, 2005, or Hilber and Vermeulen, forthcoming). This reduces the rate of new building and so over time the size of the stock of housing relative to demand within the market. This, all else equal, increases prices and thus the opportunity cost of keeping housing vacant. The effect of this is unambiguously to *reduce* vacancy rates. It will also be likely to increase price volatility.

[3] It might be argued that Travel to Work Areas (TTWA) approximate more closely to spatial housing markets but as our results demonstrate the geographical extent of both housing and labour markets is jointly determined. Planning policy is implemented by the local jurisdiction, the LA, so it is only by using these as our units of analysis that the relationship between restrictiveness and commuting distances can be revealed. Not only do TTWAs not correspond to any political jurisdiction but their boundaries are partially determined by the policy actions of their constituent jurisdictions.

However, more restrictive planning policies will also change the bundle of attributes on offer and, other things equal, slow the rate of adaptation of housing characteristics to changes in the structure of demand with respect to them – the mismatch effect. The latter effect is expected to *increase* vacancy rates. This will come about via two separate forces, one working on the characteristics of new build and the other on the adaptation of the characteristics of the existing stock of houses.

The first force may imply that new build houses become smaller, more distant from jobs, relatively more concentrated and are more likely to be in the form of flats or terraced houses, because there is less land available for dwellings. The second force arises because the structure of demand for housing characteristics changes over time and to accommodate this, the characteristics of the existing stock of housing needs to be constantly adjusted. For example, as entry to the best state schools in Britain has become more determined by the exact location of houses and the relative standing of different schools has changed, so people seeking to 'buy' entry to better state schools will want more bedroom space in the best schools' catchment areas. Another example is that as more cars have been bought (car ownership has increased 13-fold since the current form of land use planning in England was introduced in 1947 and doubled since our vacancy data starts in 1981, Department for Transport, 2013), the demand for garages and off street parking has increased. Such examples of ways in which the demand for housing attributes changes over time could be increased almost indefinitely.[4] However, what it means is that if the supply and demand for the structural characteristics of housing are to be efficiently matched to each other, there will need to be constant adaptation of the characteristics of the existing stock of houses. So LAs that are more restrictive in their application of development control will slow the adaptation of the existing stock to (changes in) the structure of demand for housing attributes.

Over time, in more restrictive LAs the characteristics of new and existing housing available will be less adapted to preferences of households. Hence, other things equal, if people have a strong idiosyncratic preference for locations and house type (e.g. a double-earner household with children that needs at least two-bedrooms and garden space), they will spend more time searching for housing that matches their preferences. When households live in a less restrictive housing market while searching in the more restrictive local market, this will imply a decrease in the vacancy rate in the former and an increase in the latter housing market, ceteris paribus.[5] In other words, given idiosyncratic preferences, households stay longer in the 'wrong' places. Households may also be induced to decrease the real costs of their search by living in suboptimal temporary accommodation such as caravans, mobile homes or their parents' spare rooms. This allows them to spend longer searching and to search more efficiently, being on the spot, but does not necessarily cause higher vacancies. It will have

[4] Other examples might be homes adapted for disabled people, or two-bedroom family homes. The relatively more scarce such houses with specific desirable characteristics are then the longer households will have to spend searching for their ideal house

[5] In Web Appendix 1 we demonstrate this in a standard search model setting, building on the seminal paper by Wheaton (1990). Using numerical simulations, we formally demonstrate that under realistic parameter assumptions an increase in the (relative) regulatory restrictiveness in a particular market increases the local vacancy rate in that market (and lowers it in the comparably less restrictive market), even with perfectly inelastic total demand for housing.

two opposing effects: on the one hand it will increase vacancies in housing because some people are housed in the temporary dwellings; but on the other, since it will improve search efficiency, it will reduce vacancies. It is an adaptive response to the relative scarcity of affordable housing matching the preferences. In Appendix 2 we provide some evidence on this adaptation. We do find that greater local restrictiveness leads to a higher share of non-permanent homes but the results are not particularly statistically strong. This may be because the measure of temporary housing is noisy.

We discuss one more obvious measure of mismatch, identified in the Introduction, in Section 3.5: commuting distances from the workplace in the LA. We do indeed find that for workers in more restrictive LAs commuting distances increase significantly. This result is consistent with house hunters finding it more difficult to match their preferences in more restrictive local housing markets so becoming 'mismatched' locationally. This has interesting implications both for the boundaries of local labour markets – they appear to be determined not just by transport costs but by local planning policies and how these affect both the total supply of housing and the supply of individual housing characteristics. But it also implies a net welfare loss, not previously observed, from more restrictive local planning. If, all else equal, commuting distances increase then there will be a net welfare loss. We return to this in section 3.5 and the conclusions.

The well-documented fact that tighter local regulation leads to higher prices is indicative that the opportunity cost effect may be important in determining local vacancy rates. However, we lack evidence on the importance of the offsetting mismatch effect. Thus the net effect of local regulatory restrictiveness on local vacancy rates is ambiguous. The empirical analysis that follows aims to identify this *net* effect while eliminating alternative explanations.

3. Empirical Analysis

3.1 Data and descriptive statistics

Our data come from several sources. The vacancy rates are from the UK Census for the years 1981, 1991, 2001 and 2011. For the first three Census years we have information on the number of vacant dwellings and we are able to distinguish between primary dwellings and second homes.[6] For 2011 the Census reported only information on the number of all unoccupied dwellings *including second homes*. Hence we use information on the number of second homes in 2001 and assume that the share of second homes remained constant between 2001 and 2011 to estimate the number of vacant dwellings for 2011. In a sensitivity check we use an alternative dataset for vacancy rates (available for 2001 and 2011 only) to test whether our findings are sensitive to this adjustment. The latter dataset is provided by the Department

[6] The Census uses the term 'household space', which is a space taken by one household, including that of just one person. Almost no household shares facilities like bathrooms (less than 0.1 percent), implying that the number of (vacant) household spaces is essentially the same as the number of (vacant) dwellings. Hence, in what follows, we will refer to dwellings as household spaces.

of Communities and Local Government (DCLG). The primary source is the Local Authority (LA) returns for the Council Tax.[7]

We obtain our measures of regulatory restrictiveness from the Planning Statistics group at the DCLG. Following the literature, our key measure is the refusal rate for major residential projects available for each LA on an annual basis. The refusal rate for 'major' projects is defined as the share of applications for residential developments of ten or more dwellings that is refused by an LA in any year during the process of 'development control'. We calculate the refusal rate for each LA using data on all applications and refused applications for the Census year itself plus the two years preceding it.[8]

As a proxy for local (housing) demand we use LA-level male weekly earnings for the period from 1981 to 2011, following Hilber and Vermeulen (forthcoming). Our earnings data come from the Annual Survey of Hours and Earnings (ASHE) for the period from 1997 to 2011 and from the New Earnings Survey (NES) for the period from 1981 to 1996. We obtained the ASHE data at the LA-level but the NES data for earlier years are only available at the county and London borough level. We then geographically matched all earnings data to the LA-level and deflated the nominal earnings figures by the Retail Price Index to obtain real earnings. For more details on the data and procedures used, we refer to Hilber and Vermeulen (forthcoming).

A number of other factors may influence vacancy rates, in particular housing tenure, demographics and socio economic characteristics. We obtain these control variables from the Population Censuses. Our list of controls includes the local homeownership rate. Homeowners tend to move less often than renters, and this is likely to be reflected in higher vacancy rates for rental housing. We also control for the share of council housing. Because rents of council houses are usually below market value, there are waiting lists for council houses. This is likely to imply a shorter duration of vacancies (Pawson and Kintrea, 2002). However, this effect could be offset if councils have less efficient housing management.

The Population Censuses also provide data on the share of people between 30 and 64 and the share of elderly, 65 and over. Young people may be more flexible in their housing choices than older people, and they may be less selective because they are more income constrained or have lower search costs (perhaps because of lower opportunity costs of time) leading to lower vacancy rates in LAs where there are proportionately more young adults. On the other hand, younger people tend to have a higher mobility rate, leading to higher vacancy rates. The mortality rate is of course highly correlated to the share of elderly. Death frequently implies that houses become vacant and, moreover, because of probate and perhaps other reasons (the new owner may not be a local resident or the house has suffered a period of neglect so is more likely to need refurbishment) houses that become vacant on the death of their owner are likely to remain vacant for longer. Other control variables derived from the

[7] The cross-sectional correlation between the 2001 Census and the DCLG data is 0.68, indicating that there are non-trivial differences in the measurement of vacant dwellings arising from the different methodologies. As is discussed later, our key findings are very similar when we use this alternative measure for vacancy rates.

[8] In a web appendix we also use additional information on the refusal rate of minor projects and show that our results are robust when we include this additional information.

Population Censuses are the share unemployed, the share of highly educated, and the share of residents with permanent illnesses.

Our instrumental variable strategy employs information on the political composition of the local council and local vote shares. We obtained the local election data from various sources: (*i)* the British Local Election Database (1889-2003) compiled by Rallings and Thrasher (2004), (*ii*) the Local Election Handbooks (1999 to 2008), (*iii*) the Local Elections Archive Project (LEAP) (2006 to 2010) and (*iv*) the BBC (2009 to 2011). We do not have data on local elections for four local authorities, so these are excluded from the analysis, leaving us with a regression sample of 350 LAs and 4 Census years (1981, 1991, 2001 and 2011).[9]

Since it might be argued that turnout is unrepresentatively low at local elections in the sensitivity analysis, we also use data on general elections, by matching each Census year to the nearest general election year (i.e., 1983, 1992, 2001 and 2010). The LA-level share of votes for the Labour party in the general elections is derived from the British Election Studies Information System. For more information on the election data, we refer to the Data Appendix.

We also gather data on house prices from the Land Registry (1995-2011) and the Council of Mortgage Lenders (CML) (1974-1995). We do so by taking account of the composition of sales in terms of housing types by adopting a mix-adjustment approach (see Wall, 1998). The real price index is obtained by again deflating the nominal series with the Retail Price Index. We then use the price index to create a measure of local price volatility; again for more information see Hilber and Vermeulen (forthcoming). Table 1 presents the descriptive statistics. It can be seen that the average overall vacancy rate is about 4 percent. The vacancy rate in 2011 was 3.6 percent. This is only slightly lower than in the United States, where it was 4.5 percent in 2012.[10] This might seem surprising when one takes into account the enormous excess supply of housing in the wake of the Great Recession that made housing extremely affordable in the US. A striking feature of vacancy rates in England is that the average vacancy rate is quite stable over time (Figure 1). In Figure 2 we plot the cross-sectional relationship between the vacancy rate and house prices. Vacancy rates are somewhat lower in areas with high prices (ρ= –0.246), consistent with the opportunity cost argument discussed above. There is little response to the housing market cycle; the correlation between the change in the vacancy rate and the change in house prices is very low with ρ= –0.069.

We map the average local vacancy rates over the sample period in Figure 3. There is meaningful variation in vacancy rates over space. Vacancy rates are generally higher in the less prosperous north. Cities like Liverpool and Bradford, which respectively relied on

[9] Since in our empirical analysis we first difference the Census data to account for time-invariant unobservable characteristics, we end up with $350 \times (4-1) = 1050$ observations. We note that some LAs have been amalgamated in 2011, reducing the total number of LAs to 326. To achieve consistency in our analysis over time we geographically match the 2011 LA information to 2001 LA boundaries with the help of official 'lookup tables'. In a robustness check we exclude those LAs that were affected by amalgamation. Results are very similar.

[10] Owner-occupied housing has a vacancy rate of 2.2 percent and rental housing about 8.8 percent. As the share of owner-occupied housing in the United States is about 0.65, the overall vacancy rate is 4.5 percent.

traditional port and port-related manufacturing or textiles, experienced decline from the 1950s. Apart from high unemployment and lower earnings there was outward migration tending to generate a more obsolete housing stock and higher housing vacancy rates. Also in areas where mining was historically important (in County Durham and Lancashire for example), vacancy rates tend to be higher. We implicitly control for these geographical differences in the industry composition by first differencing our empirical specification, thus capturing all time-invariant characteristics that vary over space. The inclusion of the first difference in local unemployment rate as a further control should effectively control for any relevant influence of changes in industrial structure on housing vacancy rates.

Refusal rates over the last 30 years have been clearly highest in the Greater London Area and in the south of England and lowest in the north of the country (Figure 4). The south of England has not only been economically considerably more successful than northern regions over the period, but it has (perhaps relatedly) had much tighter planning restrictiveness. This – despite strong housing demand – has constrained the growth of housing supply in southern England relative to the north.

3.2 Econometric framework and identification

We aim to test the impact of housing supply restrictions (as captured by the refusal rates of major projects) on vacancy rates. Let $v_{\ell,t}$ be the vacancy rate in LA ℓ in year t. $r_{\ell,t-2}$ is the refusal rate, where the refusal rate is calculated using all applications and refused applications in years $t-2, t-1$ and t. We use data up to two years before and including the year of observation to avoid random yearly fluctuations and the fact that some local authorities receive no or very few applications in a particular year. θ_t are year fixed effects that capture any aggregate economic shocks and also any policy changes at the national level that might affect vacancy rates. Then:

$$v_{\ell,t} = \alpha r_{\ell,t-2} + \theta_t + \epsilon_{\ell,t}\ , \tag{1}$$

where α is the parameter of interest and $\epsilon_{\ell,t}$ is an independently and identically distributed error term. Policy makers expect that $\alpha < 0$, implying that supply restrictions lead to a lower vacancy rate. The problem with estimating this specification using OLS is that there are potentially important endogeneity concerns with respect to $r_{\ell,t}$. First, there may be several omitted variables that have a joint impact on regulation and vacancy rates. For example, areas with more demand (higher earnings) are argued to have lower vacancy rates and more stringent planning (Hilber and Robert-Nicoud, 2013). Another concern is that due to durable housing, the north of England with its declining industries can be expected to have higher vacancy rates. It is also observed that these areas are less restrictive, so there may be spurious correlation. This may lead to a (strong) downward bias of the coefficient α. A second source of bias is that if developers know that a particular LA is more restrictive and so more likely to reject applications, they will be less likely to apply in the first place because applications cost significant resources. This implies a measurement error in the regulatory restrictiveness measure. A third concern is that vacancy rates also influence regulatory restrictiveness (reverse causality). When policy makers observe a high vacancy rate, they may become more reluctant to permit new development.

To partially address the first source of endogeneity, we estimate a first-difference equation, so that we can control for all time-invariant unobserved factors. Hence:

(2) $$\Delta v_{\ell,t} = \alpha \Delta r_{\ell,t-2} + \theta_t + \Delta \epsilon_{\ell,t}\ ,$$

where Δ denotes the change.[11]

This specification only partly addresses the first endogeneity concern because there might still be correlation with unobserved shocks. For example, in locations with increasing demand, house prices and regulatory restrictiveness may increase simultaneously. Anecdotal evidence suggests that in England regulatory restrictiveness is strongly pro-cyclical. In times of high demand, planners reject more proposals in attractive areas, perhaps to avoid 'what they perceive as a threatened 'oversupply' and perhaps because the system cannot cope with the workload. Because housing supply takes time to adjust, this will lead to lower local vacancy rates during boom periods. This again implies that α is likely strongly downward biased if we estimate (2) by OLS.

We therefore have to find an instrumental variable to identify refusal rates that is uncorrelated with local unobserved shocks. In the spirit of Bertrand and Kramarz (2002) we use the political make-up of local councils to construct an instrument for the refusal rate. Specifically, our instrument is the change in the number of seats for the Labour party between election years close to the Census years. Traditionally Labour voters and politicians have been less opposed to new residential construction than their Conservative counterparts. Labour councillors typically represent a part of the population that has less housing equity and so is less subject to NIMBY pressures aiming to protect house values. Labour councillors are also likely to be more interested in the job generating effects of construction. Thus, we can expect that an increase in the share of Labour seats may induce LAs to become less restrictive, yet, a change in the share of Labour seats should not directly affect the local vacancy rate other than through any effects it has on planning restrictiveness. So we estimate:

(3) $$\Delta v_{\ell,t} = \alpha \boldsymbol{\Delta r_{\ell,t-2}} + \theta_t + \Delta \epsilon_{\ell,t},$$

(3.1) $$\boldsymbol{\Delta r_{\ell,t-2}} = \tilde{\alpha} \Delta s_{\ell,t-2} + \tilde{\theta}_t + \Delta \tilde{\epsilon}_{\ell,t},$$

where **bold** indicates that changes in the regulatory constraints measure $\Delta r_{\ell,t-2}$ are instrumented by changes in the share Labour seats $\Delta s_{\ell t}$ and the $\sim$ refers to first stage parameters.

One might still be concerned that our instruments are correlated with $\Delta \epsilon_{\ell,t}$, so the next step is to include local authority area fixed effects η_ℓ:

(4) $$\Delta v_{\ell,t} = \alpha \boldsymbol{\Delta r_{\ell,t-2}} + \eta_\ell + \theta_t + \Delta \epsilon_{\ell,t},$$

(4.1) $$\boldsymbol{\Delta r_{\ell,t-2}} = \tilde{\alpha} \Delta s_{\ell,t-2} + \tilde{\eta}_\ell + \tilde{\theta}_t + \Delta \tilde{\epsilon}_{\ell,t}.$$

[11] One might also use a fixed effects approach. We test the robustness of the results using a fixed effects approach in Web Appendix 2 and show that results are very similar.

By including local authority area fixed effects η_ℓ, we control for all linear trends caused by unobservable factors, which increases the likelihood that changes in the instruments are uncorrelated with $\Delta\epsilon_{\ell,t}$.

If the instruments are valid (so uncorrelated with omitted variables and therefore the error term), adding additional control variables, should not influence the parameter of interest α, but also should not have an impact on the first-stage coefficients of the instrument. To test this, we include other, potentially endogenous, control variables, like changes in the demographic composition:

(5) $$\Delta v_{\ell,t} = \alpha\boldsymbol{\Delta r_{\ell,t-2}} + \beta\Delta x_{\ell,t} + \eta_\ell + \theta_t + \Delta\epsilon_{\ell,t},$$

(5.1) $$\boldsymbol{\Delta r_{\ell,t-2}} = \tilde{\alpha}\Delta s_{\ell,t-2} + \tilde{\beta}\Delta x_{\ell,t} + \tilde{\eta}_\ell + \tilde{\theta}_t + \Delta\tilde{\epsilon}_{\ell,t},$$

where $\Delta x_{\ell,t}$ is a vector of changes in the control variables. One of our control variables is the change in log local average earnings as a proxy for local demand. One might be particularly concerned about the endogeneity of earnings and we are also interested in the impact of this variable on local vacancy rates. Thus, in a robustness check, following Hilber and Vermeulen (forthcoming), we instrument for this variable using a measure that captures local demand shocks. We do not include local house prices as a control since, as we discuss in Section 2, we would expect that regulatory restrictiveness influences vacancy rates in part through house prices. Moreover, house prices and vacancy rates are jointly determined by restrictions.

The main objection to the validity of the change in the share Labour seats-instrument is that it may be correlated with (potentially non-linear) unobserved trends. For example, some local housing markets in the Greater London Area have experienced a substantial inflow of wealthy residents during the last two decades, leading to changes in the demographic composition of the local market and therefore also to changes in voting behaviour. We thus control for a flexible function of local vote shares of the previous local election, identifying regulatory restrictiveness from the random component generated by the particular feature of the British 'first past the post' voting system: seats allocated to parties are very seldom proportional to the number of votes. For example the Labour party may only receive 35 percent of the votes but since other parties also win significant vote shares, in addition to the Conservative party, Labour may still end up with a majority of seats. So what we use to identify regulatory restrictiveness is the number of seats that Labour won (or lost) beyond their vote share. While Labour's local vote share may be correlated with various demographic and socio-economic characteristics of the constituency, holding local vote shares constant, seats won (or lost) above and beyond should be uncorrelated with the error term. We can express our final estimating (base) equation as:

(6) $$\Delta v_{\ell,t} = \alpha\boldsymbol{\Delta r_{\ell,t-2}} + \beta\Delta x_{\ell,t} + \Omega(\Delta\pi_{\ell,t}) + \eta_\ell + \theta_t + \Delta\epsilon_{\ell,t},$$

(6.1) $$\boldsymbol{\Delta r_{\ell,t-2}} = \tilde{\alpha}\Delta s_{\ell,t-2} + \tilde{\beta}\Delta x_{\ell,t} + \tilde{\Omega}(\Delta\pi_{\ell,t}) + \tilde{\eta}_\ell + \tilde{\theta}_t + \Delta\tilde{\epsilon}_{\ell,t},,$$

where $\pi_{\ell,t}$ is the share of Labour votes in the closest previous local elections, and

$$(6.2) \qquad \Omega(\Delta\pi_{\ell,t}) = \sum_{n=1}^{N} \gamma_n \Delta(\pi_{\ell,t}^n) \quad \text{and} \quad \widetilde{\Omega}(\Delta\pi_{\ell,t}) = \sum_{n=1}^{N} \tilde{\gamma}_n \Delta(\pi_{\ell,t}^n).$$

Hence, $\Omega(\cdot)$ and $\widetilde{\Omega}(\cdot)$ are N^{th} order polynomials of local vote shares $\pi_{\ell,t}$ and γ_n and $\tilde{\gamma}_n$ are parameters to be estimated.

3.3 Baseline results

We start by ignoring any potential endogeneity issues and simply regress the vacancy rate on the refusal rate of major residential projects (equation 1). In Figure 5 we plot the cross-sectional relationship between the major refusal rate and the vacancy rate. We can see that refusal rates are negatively related to vacancy rates. The regression line implies that a one standard deviation increase in refusal rates is associated with a 0.23 percentage point decrease in the vacancy rate (s.e. 0.040). This naïve correlation provides 'common sense' evidence supporting the view that vacant houses can be 'regulated away'. However, the quantitative impact is not very large.

Table 2 reports estimates for equations (2) to (6). In the cases of equations (3) to (6) these are the second stage results of our IV-estimates. In column (1) we regress *the change* in the vacancy rate on *the change* in the refusal rate still ignoring potential endogeneity issues (equation 2). We first difference controls to offset for any time invariant omitted characteristics such as differences in income levels across local authorities. We see that even without instrumenting for the refusal rate or adding control variables, the relationship between (the changes in) the planning restrictiveness and (the changes in) the vacancy rate is no longer negative and statistically significant.

However, because of the endogeneity concerns discussed, the coefficient on the refusal rate cannot be interpreted as a causal effect. As a first step to addressing this issue we include local authority fixed effects in column (2). The coefficient on the change in major refusal rate now becomes positive and statistically significant at the 10 percent level. In column (3) of Table 2 we add further control variables as discussed in Section 3.1 above. The estimated coefficient for the change in the major refusal rate is hardly affected, although it is not statistically significant at conventional levels anymore. The control variables often have a statistically significant impact on the change in the vacancy rate with the anticipated sign. For example, areas with an increasing share of elderly people or of council housing experience an increase in the vacancy rate. Also, areas with an increasing unemployment rate, from which people may have been tending to move away, experience an increase in the vacancy rate. In areas with a rising share of highly educated people, vacancies tend to decrease.

Still, however, regulatory restrictiveness is likely measured with error (because developers may not apply in the first place in more restrictive places). It may also be correlated with unobserved shocks. Moreover, we should address the potential reverse causality issue that higher vacancy rates may induce policy makers to be more restrictive. We therefore instrument for the change in the major refusal rate with the change in the share Labour seats in column (4). This specification corresponds to equation (3) above.

Kleibergen-Paap *F*-statistics indicate that weak identification of regulatory restrictiveness is not an issue. The results suggest that a standard deviation increase in the major refusal rate leads to an increase in the vacancy rate of 0.82 percentage points. As noted in the previous subsection, one objection to the instrument is that it may be correlated with unobserved characteristics of the area. To control for this, we include LA fixed effects in column (5) – corresponding to equation (4). The coefficient on the refusal rate is hardly changed and remains statistically significant at the five percent level. Column (6), corresponding to equation (5), includes the same range of control variables as in column (3). This makes almost no difference to the estimated coefficient of primary interest.

One might still be worried that changes in the share of Labour seats are correlated with unobservable shocks (e.g. gentrification) that simultaneously have an impact on voting behaviour and vacancy rates. So in column (7) we estimate our final model (6). That is, we additionally include a flexible function of changes in the Labour *votes* in local elections, approximated by a fifth-order polynomial to isolate the impact of voting behaviour caused by the demographic and socio-economic composition of the LA from political power (measured by *seats*). In the sensitivity checks, discussed below, we report results for different orders of polynomials. Reassuringly, the estimated effect of regulatory restrictiveness in column (7) is very similar to the previous specifications. The instrument is somewhat less strong (with a Kleibergen-Paap *F*-statistic of 8.2). Still, we find a positive and economically meaningful effect of regulatory restrictiveness on the vacancy rate: a one standard deviation increase in the major refusal rate increases the vacancy rate by 0.90 percentage points. Due to the correlation between changes in the Labour vote shares and changes in the share of Labour seats, it is no surprise that the coefficient is now only statistically significant at the 10 percent level.[12]

In Table 3 we report the corresponding first-stage estimates: a standard deviation increase in the share of Labour seats leads to a decrease in the major refusal rate of 0.26-0.34 standard deviations. It is notable that the first-stage coefficients of the change in the share Labour seats instrument are highly statistically significant and are hardly affected by the inclusion of local authority fixed effects and other control variables. If we include vote share controls, the coefficient on change in Labour seats becomes slightly lower, but it is still statistically significant at the 5 percent level.

3.4 Sensitivity analysis

In this subsection, we conduct an extensive set of sensitivity analyses. The first concern we address is that the data on vacant housing may be measured with error, because in 2011 the Census also included second homes. We therefore use another data source for 2001 and 2011 from the Department of Communities and Local Government (DCLG - only available for a shorter time period). Because we do not have clear priors which data source provides a better estimate of 'real' vacancies in 2001, we calculate the average vacancy rate using both data sources for the year. We then estimate the same models as in Table 2. The second-stage

[12] The correlation between the share Labour votes and the share Labour seats is 0.88. However, the correlation between the change in share Labour votes and the change in share Labour seats is much lower (ρ= 0.481).

results are reported in Table 4. Column (1) shows that even the bivariate specification suggests a positive and significant correlation between changes in regulatory restrictiveness and changes in vacancy rates. This also holds if we include local authority fixed effects in column (2) and control variables in column (3). In column (4), we instrument the change in major refusal rate with the change in share of Labour seats. The coefficient is almost identical to previous results: a one standard deviation increase in the major refusal rate increases the vacancy rate by 1.0 percentage points. This result is hardly affected if we include local authority fixed effects in column (5) and control variables in column (6). In column (7), we finally control flexibly for the change in local vote shares. In the last specification the coefficient on the change in the major refusal rate variable becomes slightly but not statistically significantly smaller. It is still statistically significant at the 10 percent level. Note that the first-stage estimates are identical to the ones presented in Table 3. The results seem highly reassuring and strongly indicate that the potential measurement error in the Census data is not influencing our results.

In Table 5 we further test the robustness of our results to the potential measurement problem by excluding 2011 from the analysis. We report OLS and second stage results in Panel A. In columns (1) to (3), where we do not address endogeneity concerns, changes in the major refusal rate are positively associated with changes in the vacancy rate, although the effect is not statistically significant in column (3). In column (4), we instrument for the change in the major refusal rate. Again, the coefficient is strongly positive, but quite imprecisely estimated. The same holds for the remaining models; due to weak identification (see Panel B), the estimated effects are rather imprecisely estimated. Nevertheless, they seem to point towards a positive and economically meaningful effect of regulatory restrictiveness on vacancy rates, in line with the previous results.

Recall that in our preferred specification (column (7) of Table 2), we included a fifth-order polynomial of changes in the share Labour votes in each LA. This is to isolate the impact of potentially unobserved demographic and socio-economic variables, which may be reflected in voting behaviour, from local political power. However, the choice of the order of polynomial is somewhat arbitrary. Table 6 investigates the robustness of the results to this choice. Panel A reports second-stage results, whereas Panel B reports the corresponding first-stage results. In column (1) of Panel A we include only a linear term of change in the share Labour votes in local elections as a control. Changes in the share of Labour votes do not have a direct effect on changes in vacancy rates. The coefficient on the change in the major refusal rate variable is statistically significant at the 11 percent level. When we include a third or fourth order polynomial of change in share Labour votes, the coefficient on the change in the major refusal rate variable becomes slightly higher and statistically significant at the 10 percent level. In column (5) we include a fifth-order polynomial of the change in local Labour vote shares, but we also include a fifth-order polynomial of the change in general election vote shares. The latter might be relevant, as one could argue that results of general elections might be a better proxy for the demographic characteristics of a local authority, due to the substantially higher turnouts; on the other hand re-working the Parliamentary Constituency vote shares to generate an estimate for LAs must induce some measurement error. The point estimate is very similar to the preferred specification, but the effect is only

statistically significant at the 20 percent level, likely due to a weaker first-stage (the corresponding first-stage Kleibergen-Paap *F*-statistic is only 4.9). In Panel B, we report corresponding first-stage estimates. The instrument has a similar impact across different specifications, but becomes somewhat smaller in magnitude, once we allow for more flexibility in the vote share controls.

In the baseline results we only treat changes in the major refusal rate as endogenous. It is possible to argue that changes in earnings are also subject to endogeneity concerns. Earnings, which are a proxy for local demand for housing, are also dependent on the reaction of labour supply to changes in demand. In turn, local labour supply depends on the flexibility and adaptability of the housing stock to accommodate new workers, and may therefore depend on the vacancy rate as well (Glaeser *et al*., 2006, Saks, 2008, Hilber and Vermeulen, forthcoming). We therefore use a labour demand shock measure based on employment by industry in 1971 as an instrument for earnings. The shock predicts the level of employment in each local authority using information on national employment growth in each industry. So, we use exogenous changes (from the local perspective) in employment growth to predict total employment in each local authority in each year. For example, this measure predicts a large drop in employment in areas that were specialised in mining in 1971. For a fuller discussion see Hilber and Vermeulen (forthcoming).

Table 7 reports the results. Panel A reports second-stage results, while Panels B and C report the corresponding first-stage results. In column (1) of Panel A we only include the endogenous variables change in major refusal rate and change in earnings plus the year fixed effects. Changes in regulatory restrictiveness are still strongly positively correlated with changes in vacancy rates. The effect of changes in earnings is statistically insignificant, in line with previous results. Column (2) additionally includes LA fixed effects. The Kleibergen-Paap *F*-statistic of 5.5 now indicates fairly weak identification. Since our model is just identified, the estimated instrumented coefficients are median unbiased, yet they may be too imprecisely estimated to be useful (Angrist and Pischke, 2009). Moreover, if our results are biased, they should be biased towards the corresponding OLS estimates (see columns (2) and (3) Table 2), which in our case would imply that the estimated coefficients are themselves *under*estimates (Stock and Yogo, 2005; Murray, 2006). In any case with these caveats in mind, column (2) indicates that changes in the regulatory restrictiveness have a positive and statistically significant impact on changes in vacancy rates. Although the coefficient on the change in earnings variable now becomes positive and is much larger in magnitude than in previous specifications, it is not statistically significantly different from zero. When we include further control variables in column (3), the results are hardly affected (and coefficients remain weakly identified). In the final – most rigorous – specification, reported in column (4), we also control flexibly for vote shares. Thus we identify our endogenous variables off of random variation in the number of Labour seats won beyond the share of Labour votes. The major refusal rate has a positive and statistically significant impact (at the five percent level) on vacancy rates: a one standard deviation increase in the major refusal rate increases the vacancy rate by 1.1 percentage points, very similar to that implied by the estimates reported in Table 2.

In Panels B and Panel C we report the corresponding first-stage estimates of the models for the change in the major refusal rate and the change in earnings respectively. Changes in the share of Labour seats are a reasonably strong instrument for the change in the refusal rate. Changes in the labour demand shock measure are strongly positively correlated to changes in the refusal rate (i.e. areas that have experienced an exogenous inflow of employment have also become substantially more restrictive). In Panel C we observe that changes in the labour demand shock measure are also positively correlated to changes in earnings, as anticipated.

In Web Appendix 2 we conduct a series of additional sensitivity analyses. We show that our results are robust to (*i*) excluding the Greater London area, (*ii*) using the refusal rate on all projects – not just major ones – and (*iii*) using a fixed effects, rather than a first-differencing approach.

To conclude, the various robustness checks all deliver very similar estimated effects of restrictiveness on vacancy rates in terms of magnitude compared to the baseline models. This provides additional support for the proposition that increased regulatory restrictiveness causes higher vacancy rates. We note that the results are not always statistically strong in the more comprehensive specifications. This appears to be mainly due to weak(er) identification.

3.5 Direct measures of mismatch and regulatory restrictiveness

In Section 2 we hypothesised that a positive relationship between restrictions and vacancy rates might be explained by increased mismatch. There are few obvious measures of mismatch but for reasons discussed above we think that the 'average commuting distance from the workplace' does not only provide a useful measure but should further reveal the underlying interrelationship between housing and labour markets. One of the most important characteristics of a house is its location with respect to jobs. It seems reasonable therefore that the average commuting distance from the workplace should capture mismatch in this dimension of housing characteristics for any given housing market. In principle, households have a preference to live close to their workplaces. If regulatory restrictions make it more difficult for people to find a home 'matched' to their preferences on other characteristics close to work, they have to search for properties further away. This adaptation of search behaviour implies, other things equal, vacancies will tend to be higher in the more restrictive LAs and lower in neighbouring, less restrictive ones, as workers become more mismatched locationally.

We gather data on the average commuting distance from the workplace for all the Census years. The data provide us with the share of people per commuting distance band (0-2 km, 2-5 km, etc.). We then calculate the average commuting distance by taking the midpoint of each category and weighting it by the number of persons in each category.

Panel A in Table 8 reports the regression results of the log of average workplace commuting distance on the major refusal rate, where we follow the same approach as in Table 2. The results in the first column seem to suggest that commuting distance is not influenced by regulatory restrictiveness. When we include LA fixed effects and demographic control variables, the results are still statistically insignificant. This is not too surprising as the refusal rate is highly endogenous and correlated to other factors that might explain commuting

distances. For example, places that have become denser tend to have become more restrictive, but denser places also might have shorter commutes because jobs and households are located closer to each other.

We therefore control for other factors that might be correlated with the refusal rate by instrumenting for the change in the refusal rate with the change in the share of labour seats (as in Table 2). This reveals a positive and significant effect - see Panel A of Table 8, column (4). As local restrictiveness in the LA in which a worker is employed increases so does the average commuting distance: a one standard deviation increase in the major refusal rate increases the commuting distance by 8.5 percent, a non-negligible effect. The effect becomes somewhat smaller (5.8 percent) when we include in column (5) local authority fixed effects. The effect continues to be essentially the same when we add further control variables in column (6) and a flexible function of the share of labour votes in column (7). In the last column, however, the effect is somewhat imprecisely estimated and only statistically significant at the 14 percent level.[13]

As noted in Section 2 we also experimented with other measures potentially capturing mismatch. In Appendix 2 we report results for the share of non-permanent dwellings. Our underlying explanation for why this measure should proxy for housing market mismatch is that in more restrictive markets there is an incentive to accept even less optimally matched housing characteristics in the short term in order to intensify and increase the efficiency of search. Living in temporary accommodation has a low switching cost associated with it and is a cheaper strategy than buying a suboptimal place to live and then reselling it when a more suitable house is found. The ease with which search can be undertaken and its effectiveness will increase if the house-hunter can be physically present in the local market (Ha and Hilber, 2013). Moreover since the chances of finding a better match in the housing market will improve with length of time spent searching then there will be a payoff to having temporary accommodation available for searchers in markets where matching is more difficult. Thus in more restrictive LAs, other things equal, matching is more difficult and the share of temporary dwellings is greater. Although by improving the efficiency of search, temporary housing may itself reduce vacancies of permanent dwellings, it is such a relatively sub-optimal form of housing we would expect house-hunters to resort to it only when there is extreme difficulty in matching their preferences to available housing supply. So we would expect that net the share of temporary housing would be positively correlated with local restrictiveness. As reported in Table A2 in Appendix 2 we find weak, but consistent, evidence that more restrictive markets will lead to a higher rate of non-permanent homes. Given that the rate of non-permanent homes is a somewhat noisy and indirect proxy for mismatch, it may not be too surprising that the results are not statistically highly significant.

[13] One may argue that earnings are endogenous, leading to biased results. However, if we instrument for earnings with a labour demand shock variable (as in Panel A of Table 7), the results are very similar. The point estimates related to the major refusal rate are around 7 percent, while the standard errors are even somewhat lower.

3.6 Are there other potential explanations for the increase in housing vacancies with restrictiveness?

There may also be other explanations for the positive relationship between restrictions and vacancy rates. In particular, one might expect that greater price volatility is associated with higher vacancy rates. This is because price volatility might create a (real) 'option to wait' (McDonald and Siegel, 1986; Grenadier, 1995 and 1996). The greater the uncertainty (price volatility) the more valuable is a property owner's option to delay selling and renting out the property. Thus, in times of greater uncertainty, owners will be more reluctant to sell or rent out their property. Especially in markets with lengthy leases – such as office markets – this can generate high and persistent ("sticky") vacancy rates; landlords are better off keeping their units empty.

We would expect however the real options argument to be less important in the British residential property markets compared to office markets since demand volatility tends to be lower (people have to live somewhere) and, for rentals the lease length is typically quite short (a year or even less).

The real options argument, nevertheless, could be relevant since tight regulation (more inelastic supply) amplifies demand shocks, so we would expect price volatility to respond more strongly to demand volatility in places with tighter regulatory constraints. The results reported in Panel B of Table 8 show that at least according to the instrumental variable specifications (columns 4 to 7) the sensitivity of price volatility with respect to earnings volatility increases with regulatory restrictiveness. This provides some evidence that regulatory constraints *potentially* can increase vacancy rates by increasing price volatility. So the value of the real option to keep properties empty could in principle be an explanation for the positive relationship between regulatory restrictiveness and the local vacancy rate.

In Appendix 2 we also estimate regressions where we directly control for commuting distance from the workplace and price volatility. In Table A3 we show that the coefficient on the refusal rate variable decreases by about 15 percent and ceases to be statistically significant at conventional levels once we control for commuting time (column 1). This suggests that at least a part of the positive effect of a change in regulatory restrictiveness on the change in vacancy rates is driven by mismatch in the housing market, as proxied by commuting time. When we instead control for price volatility (column 2), interestingly, the effect of the refusal rate on vacancy rates increases in magnitude and is statistically significant. This finding could be interpreted as suggesting that the real options argument does not play an important role in explaining the positive link between regulatory restrictiveness and vacancy rates. When we include both variables – commuting time and price volatility – in column (3), the coefficient on the instrumented refusal rate variable is again very similar to our base specification. When we finally additionally include the change in the share of non-permanent homes to our specification (column 4), the coefficient is again somewhat reduced providing some mild additional supportive evidence for the mismatch mechanism as the primary driving force.

We should note two caveats. First, our proxies, for mismatch in particular, are at best partial and, in the case particularly of the share of temporary homes, indirect, so we would not expect these variables to fully account for the positive impact of the (change in the) refusal rate on the (change in the) vacancy rate. Second, the findings reported in Table A3 should generally be interpreted with some caution. This is because both the commuting distance and the price volatility are likely endogenous, so the reported coefficients are likely biased. Still, overall, we interpret the findings from Table A3 as indicating that our key finding of a positive impact of local regulation on local vacancy rates is largely driven by the mismatch between the preferences of the local residents and the characteristics of the available local housing stock. This finding is certainly plausible in the context of the extraordinarily rigid British planning system. Whether similar effects can be observed in other countries is an interesting question.

4. Conclusions

This is the first attempt to rigorously analyse spatial and temporal variation in housing vacancy rates. It would come as no surprise to economists to observe that in well-functioning labour markets there was unemployment. Workers search for jobs and employers seek (better) qualified workers. Attempting to regulate unemployment away makes no sense. Vacant houses are equivalent to unemployed workers yet, at least in Britain, policy does try to 'regulate' vacant homes away by using their existence to be more restrictive in the control of the supply of new homes and the structural adaptation of existing ones.

In this paper we argue that such restrictions have two main opposing effects on housing vacancies. The 'opportunity cost effect' leads to a lower vacancy rate in the more restrictive housing markets because supply constraints lead to higher prices, and thus to higher opportunity costs of keeping housing vacant. The 'mismatch effect', however, implies higher vacancy rates in the more restrictive housing markets because households will find it more difficult to match their preferences to the characteristics of the local housing supply given their budget constraints. So search becomes more prolonged and costly.

We do indeed confirm that there is a simple negative correlation between local land use planning restrictiveness and local housing vacancies. Superficially this appears to support the planners' 'common sense' that the existence of empty houses means they can – even should – plan to be more restrictive in supply. This unconditional correlation, however, is the result of a form of joint causation. When we effectively control for unobserved and unobservable characteristics at the local level by using first differencing and add year fixed effects, to control for macroeconomic shocks, the negative correlation turns positive. When we further control for local linear trends, for other potential explanatory variables and account for the endogeneity of local regulatory restrictiveness, the causal effect of restrictiveness on vacancy rates is firmly positive in both a statistically significant and economically substantial manner.

Our empirical analysis does not fully unpack the box of explanations. It is a reduced form telling us what the net impact of increased planning restrictiveness is on housing vacancies. If an LA becomes more restrictive, signalled by an increase in the rate of refusal of residential

development proposals, then all else equal vacancy rates increase in that LA. We also provide direct evidence that mismatch may be an important reason for higher vacancy rates in more restrictive housing markets: The more restrictive in planning terms a local jurisdiction is, the longer is the average commuting distance of workers within it. We subject our findings to an extensive sensitivity analysis. They survive remarkably unaltered.

Welfare implications in markets with search frictions are not easy to derive. This is because households may search less or more than would be welfare optimal (Koster and Van Ommeren, 2015). However, if search levels are close to optimal, local regulatory constraints will likely lead to lower welfare levels via increasing local vacancy rates. This is because new homes are not built and existing homes are not renovated (adapted) to reflect the current preferences for housing characteristics in places with high demand. Moreover tight regulatory constraints also prevent conversion of existing non-residential property stock to housing in these places.

It is the mismatch between the preferences of households and the housing stock on offer that leads, other things equal, to higher vacancy rates in the more regulated – typically more desirable – places. This is not to say that tighter regulatory constraints in some places necessarily increase vacancy rates at the aggregate level (for the whole country). However, even if on aggregate vacancy rates were unaffected by regulatory restrictiveness, regulatory constraints still likely cause a significant welfare loss. This is because too much housing stays empty in the most regulated, most desirable places with the strongest demand and highest valuations for living space (i.e., in the "wrong places") and people are induced to commute further, itself an additional and net welfare loss.

There are important implications for policy, particularly for the UK because of its extraordinarily restrictive planning system. Crucially planners should not allocate less land for development on the grounds that there are empty houses. Some vacancies are integral to the well-functioning of any market and, as our results show, trying to 'regulate housing vacancies away' is counterproductive. There is moreover a nice irony for advocates of the 'compact city'. The most common policy to attempt to implement this ideal is to impose growth boundaries (make land scarcer) and be more restrictive to adaptations of the existing stock or – in the US – make it more difficult to obtain zoning ordinance waivers. In summary aiming for a compact city tends make planning policy more restrictive. Our results show this will have exactly the opposite to the intended effect because average commuting distances will lengthen as residents search further afield for housing they can afford which matches their preferences.

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TABLES

TABLE 1
Descriptive statistics (repeated cross-section)

	Mean	Std. Dev.	Min	Max
Vacancy rate (in %)	3.886	1.340	0	12.06
Major refusal rate, *t-2* (in %)	27.43	14.63	0	78.57
Male weekly earnings (in £)	545.7	147.3	258.3	1,793
Share owner-occupied housing	0.669	0.112	0.0461	0.895
Share council housing	0.161	0.113	0.00416	0.820
Share age 30-65	0.449	0.0278	0.364	0.516
Share age >65	0.164	0.0364	0.0613	0.314
Share unemployed	0.0652	0.0296	0.0204	0.224
Share highly educated	0.137	0.114	0.00244	0.536
Share permanent illness	0.0348	0.0188	0.00745	0.122
Predicted employment ('labour demand shock') †	57,913	47,674	9,832	474,473
Share labour seats, *t-2*	0.309	0.269	0	0.992
Share labour voters, local elections	0.315	0.168	0	0.770
Share labour voters, general election	0.295	0.152	0.0243	0.753
Mean commuting distance from workplace (in km)	6.853	3.465	2.141	18.66
Coefficient of variation house prices, *t+3*	0.104	0.0745	0.00755	0.460
Coefficient of variation earnings, *t+3*	0.0755	0.0400	0.00685	0.674
Rate non-permanent homes (in %)	0.572	0.591	0.00670	5.862

Notes: The number of observations is 1,400, as we have 4 observations for 350 local authorities. *t-2* denotes that we include applications up to two years preceding and including the year of observation, *t+4* denotes that we include data up to four years after and including the year of observation. † Measure is based on 1971 local industry composition.

TABLE 2

Baseline results – second stage

(Dependent variable: Δ Vacancy rate)

	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS
Δ Major refusal rate, *t-2*	0.0434	0.101*	0.0801	**0.820****	**0.855****	**0.840****	**0.895***
	(0.0382)	(0.0532)	(0.0501)	**(0.417)**	**(0.405)**	**(0.354)**	**(0.530)**
Δ Earnings *(log)*			0.373			-0.0701	-0.125
			(0.677)			(0.692)	(0.719)
Δ Share owner-occupied housing			-0.129			-0.262	-0.321
			(0.356)			(0.369)	(0.418)
Δ Share council housing			0.736***			0.655***	0.631***
			(0.235)			(0.219)	(0.233)
Δ Share age 30-65			-0.0989			0.0550	0.0754
			(0.208)			(0.202)	(0.211)
Δ Share age >65			1.281***			1.432***	1.362***
			(0.317)			(0.311)	(0.319)
Δ Share unemployed			0.368***			0.270**	0.276**
			(0.138)			(0.114)	(0.136)
Δ Share highly educated			-0.493**			-0.581***	-0.617***
			(0.191)			(0.189)	(0.214)
Δ Share permanent illness			0.192**			0.151	0.123
			(0.0972)			(0.0934)	(0.0985)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.373	0.488	0.559				
Kleibergen-Paap *F*-statistic				18.52	17.49	19.01	8.151

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. In Models (4)-(7), the instrument for Δ Major refusal rate is Δ Share of labour seats in the local council. $\Omega(\cdot)$ is approximated by a fifth-order polynomial of share labour voters in local elections. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 3
Baseline results – first stage
(Dependent variable: Δ Major refusal rate, t-2)

	(4) OLS	(5) OLS	(6) OLS	(7) OLS
Δ Share labour seats, *t-2*	-0.301***	-0.322***	-0.343***	-0.264**
	(0.0700)	(0.0943)	(0.0966)	(0.114)
Δ Earnings *(log)*			0.579	0.504
			(0.461)	(0.468)
Δ Share owner-occupied housing			0.250	0.318
			(0.327)	(0.323)
Δ Share council housing			0.0880	0.113
			(0.186)	(0.188)
Δ Share age 30-65			-0.194	-0.184
			(0.150)	(0.151)
Δ Share age >65			-0.160	-0.148
			(0.220)	(0.218)
Δ Share unemployed			0.112	0.134
			(0.0880)	(0.0874)
Δ Share highly educated			0.167	0.189
			(0.171)	(0.175)
Δ Share permanent illness			0.0828	0.0673
			(0.0826)	(0.0865)
Local authority fixed effects (350)	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Δ Share labour voters local elections $\widetilde{\Omega}(\cdot)$	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050
Adj. R-squared	0.168	0.310	0.324	0.331

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. $\widetilde{\Omega}(\cdot)$ is approximated by a fifth-order polynomial of share labour voters in local elections. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 4

Results using Census data (1981-2001) and DCLG data (2001-2011)

(Dependent variable: Δ Vacancy rate)

	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS
Δ Major refusal rate, *t-2*	0.100***	0.141***	0.115**	**0.961*** **	**1.040*** **	**0.917*** **	**0.780* **
	(0.0375)	(0.0502)	(0.0472)	**(0.350)**	**(0.348)**	**(0.308)**	**(0.429)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	No	No	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.294	0.468	0.522				
Kleibergen-Paap *F*-statistic				18.56	17.61	19.09	8.212

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. In Models (4)-(7), the instrument for Δ Major refusal rate is Δ Share of labour seats in the local council. $\Omega(\cdot)$ is approximated by a fifth-order polynomial of share labour voters in local elections. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 5

Results using Census data (1981-2001, excluding 2011)

PANEL A – Second stage	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(Dependent variable: Δ Vacancy rate)	OLS	OLS	OLS	2SLS	2SLS	2SLS	2SLS
Δ Major refusal rate, *t-2*	0.123**	0.191**	0.114	**1.121**	**2.483***	**2.321***	**2.844**
	(0.0490)	(0.0896)	(0.0702)	**(0.697)**	**(1.278)**	**(1.401)**	**(2.467)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	No	No	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes
Observations	700	700	700	700	700	700	700
Adj. R-squared	0.412	0.620	0.736				
Kleibergen-Paap *F*-statistic				7.528	4.365	2.939	1.355

PANEL B – First stage	(4)	(5)	(6)	(7)
(Dependent variable: Δ Major refusal rate, t-2)	OLS	OLS	OLS	OLS
Δ Share labour seats, *t-2*	-0.243***	-0.255	-0.219	-0.163
	(0.0886)	(0.173)	(0.182)	(0.200)
Year fixed effects	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	Yes
Observations	700	700	700	700
Adj. R-squared	0.070	0.358	0.375	0.385

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. In Models (4)-(7), the instrument for Δ Major refusal rate is Δ Share of labour seats in the local council. $\Omega(\cdot)$ is approximated by a fifth-order polynomial of share labour voters in local elections. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 6
Controlling for vote shares

PANEL A – Second stage	(1)	(2)	(3)	(4)	(5)
(Dependent variable: Δ Vacancy rate)	2SLS	2SLS	2SLS	2SLS	2SLS
Δ Major refusal rate, *t-2*	**0.756**	**0.782**	**0.861***	**0.897***	**0.888**
	(0.473)	**(0.490)**	**(0.505)**	**(0.526)**	**(0.689)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	Yes	Yes	Yes	Yes	Yes
Control variables included (8)	Yes	Yes	Yes	Yes	Yes
Δ Share labour votes local elections	-0.0742	2nd order	3rd order	4th order	5th order
	(0.158)	polynomial	polynomial	polynomial	polynomial
Δ Share labour voters general elections	No	No	No	No	5th order polynomial
Observations	1,050	1,050	1,050	1,050	1,050
Kleibergen-Paap *F*-statistic	9.213	8.896	8.680	8.285	4.917

PANEL B – First stage	(1)	(2)	(3)	(4)	(5)
(Dependent variable: Δ Major refusal rate, t-2)	OLS	OLS	OLS	OLS	OLS
Δ Share labour seats, *t-2*	**-0.274****	**-0.271****	**-0.273****	**-0.267****	**-0.215***
	(0.111)	**(0.112)**	**(0.114)**	**(0.114)**	**(0.119)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	Yes	Yes	Yes	Yes	Yes
Control variables included (8)	Yes	Yes	Yes	Yes	Yes
Δ Share labour votes local elections	-0.174*	2nd order	3rd order	4th order	5th order
	(0.101)	polynomial	polynomial	polynomial	polynomial
Δ Share labour voters general elections	No	No	No	No	5th order polynomial
Observations	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.329	0.329	0.329	0.331	0.338

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. The instrument for Δ Major refusal rate, *t-2* is Δ Share labour seats, *t-2*. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 7

Instrumenting for regulatory restrictiveness and earnings

PANEL A – Second stage *(Dependent variable: Δ Vacancy rate)*	(1) 2SLS	(2) 2SLS	(3) 2SLS	(4) 2SLS
Δ Major refusal rate, *t-2*	**0.814****	**0.966****	**0.816***	**1.055****
	(0.390)	**(0.387)**	**(0.439)**	**(0.498)**
Δ Earnings *(log)*	**-0.257**	**4.469**	**6.370**	**4.149**
	(1.717)	**(5.170)**	**(5.820)**	**(5.114)**
Year fixed effects	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	Yes
Δ Share labour voters local elections Ω(·)	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050
Kleibergen-Paap *F*-statistic	9.660	5.467	3.714	3.610

PANEL B – First stage *(Dependent variable: Δ Major refusal rate, t-2)*	(1) OLS	(2) OLS	(3) OLS	(4) OLS
Δ Share labour seats, *t-2*	-0.305***	-0.297***	-0.312***	-0.224*
	(0.0702)	(0.0957)	(0.0981)	(0.116)
Δ Labour demand shock 1971 *(log)*	2.121***	6.953**	10.29***	10.12***
	(0.666)	(2.843)	(3.102)	(3.159)
Year fixed effects	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	Yes
Δ Share labour voters local elections Ω(·)	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050
Adj. R-squared	0.172	0.317	0.335	0.342
Angrist-Pischke *F*-statistic	14.12	14.17	19.00	12.24

PANEL C – First stage *(Dependent variable: Δ Earnings (log))*	(1) OLS	(2) OLS	(3) OLS	(4) OLS
Δ Share labour seats, *t-2*	0.00575	0.0119	0.00259	0.0140
	(0.00864)	(0.0117)	(0.0106)	(0.0124)
Δ Labour demand shock 1971 *(log)*	0.823***	1.011***	1.245***	1.267***
	(0.0809)	(0.283)	(0.326)	(0.339)
Year fixed effects	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	Yes
Δ Share labour voters local elections Ω(·)	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050
Adj. R-squared	0.406	0.557	0.616	0.622
Angrist-Pischke *F*-statistic	51.73	9.56	11.30	10.67

Notes: **Bold** indicates instrumented. The instruments for Δ Major refusal rate and Δ Earnings *(log)* are Δ Share of labour seats and Δ Labour demand shock 1971 *(log)*. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10.

TABLE 8

Mismatch, real options and regulatory restrictiveness

PANEL A – Second stage (Dependent variable: Δ Commuting distance from workplace (log))	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS
Δ Major refusal rate, *t-2*	-0.00274	0.00341	0.00106	**0.0847*** **	**0.0584** **	**0.0594** **	**0.0608**$^{+}$
	(0.00283)	(0.00400)	(0.00384)	**(0.0321)**	**(0.0296)**	**(0.0274)**	**(0.0407)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	No	No	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.883	0.913	0.922				
Kleibergen-Paap *F*-statistic				18.52	17.49	19.01	8.151

PANEL B – Second stage (Dependent variable: Δ Coefficient of variation of house prices, t+3)	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS
Δ Major refusal rate, *t-2*	0.00499	-0.0493***	0.00855	**-0.0891***	**-0.119** **	**-0.0726***	**-0.116***
	(0.00453)	(0.0126)	(0.00659)	**(0.0479)**	**(0.0605)**	**(0.0434)**	**(0.0679)**
Δ Major refusal rate, *t-2* × Coefficient of variation of earnings, *t+3*	-0.0373	0.00700	-0.0757	**0.802***	**1.165** **	**0.692***	**0.811***
	(0.0605)	(0.155)	(0.0858)	**(0.410)**	**(0.570)**	**(0.367)**	**(0.468)**
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	No	No	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes
Observations	1,050	1,050	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.332	0.153	0.527				
Kleibergen-Paap *F*-statistic				18.52	17.49	19.01	8.151

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. In Models (4)-(7), instruments for Δ Major refusal rate and Δ Major refusal rate × Coefficient of variation of earnings are Δ Share of labour seats and Δ Share of labour seats × Coefficient of variation of earnings. $\Omega(\cdot)$ is approximated by a fifth-order polynomial of share labour voters. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10, $^{+}$ p<0.20.

FIGURES

FIGURE 1

Vacancy rate and house prices in England between 1981 and 2011

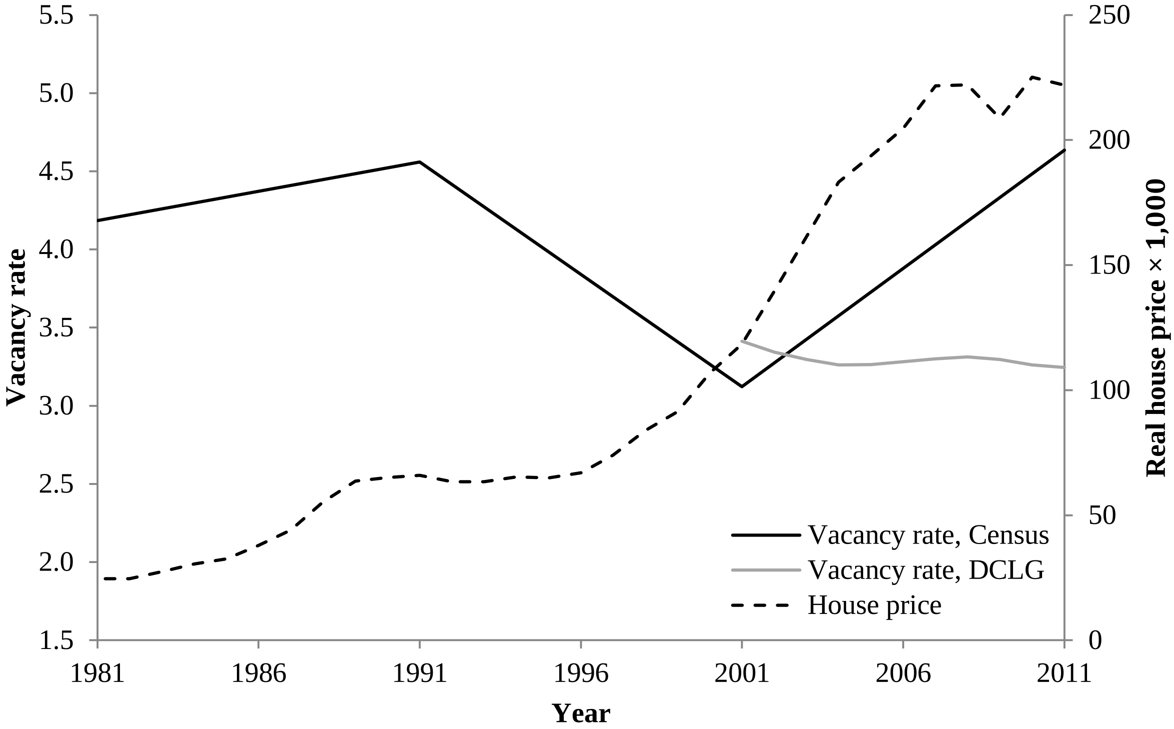


FIGURE 2

Correlation between vacancy rate and real house price

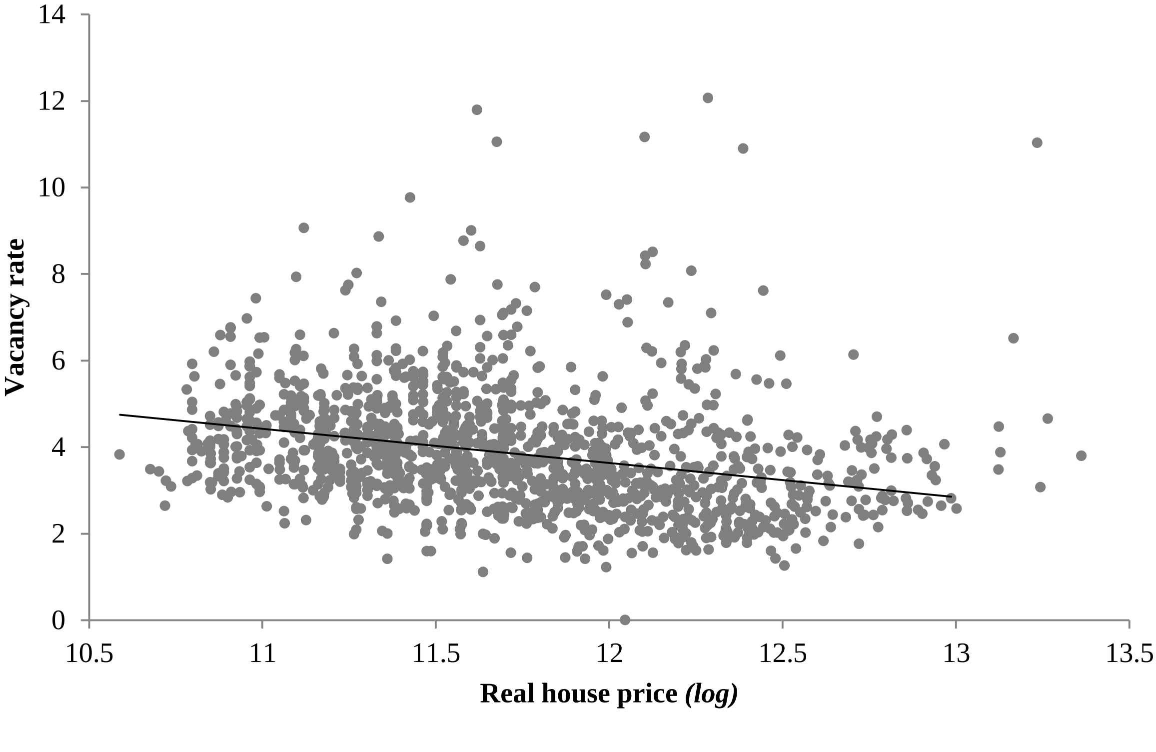


FIGURE 3
Vacancy rate across England

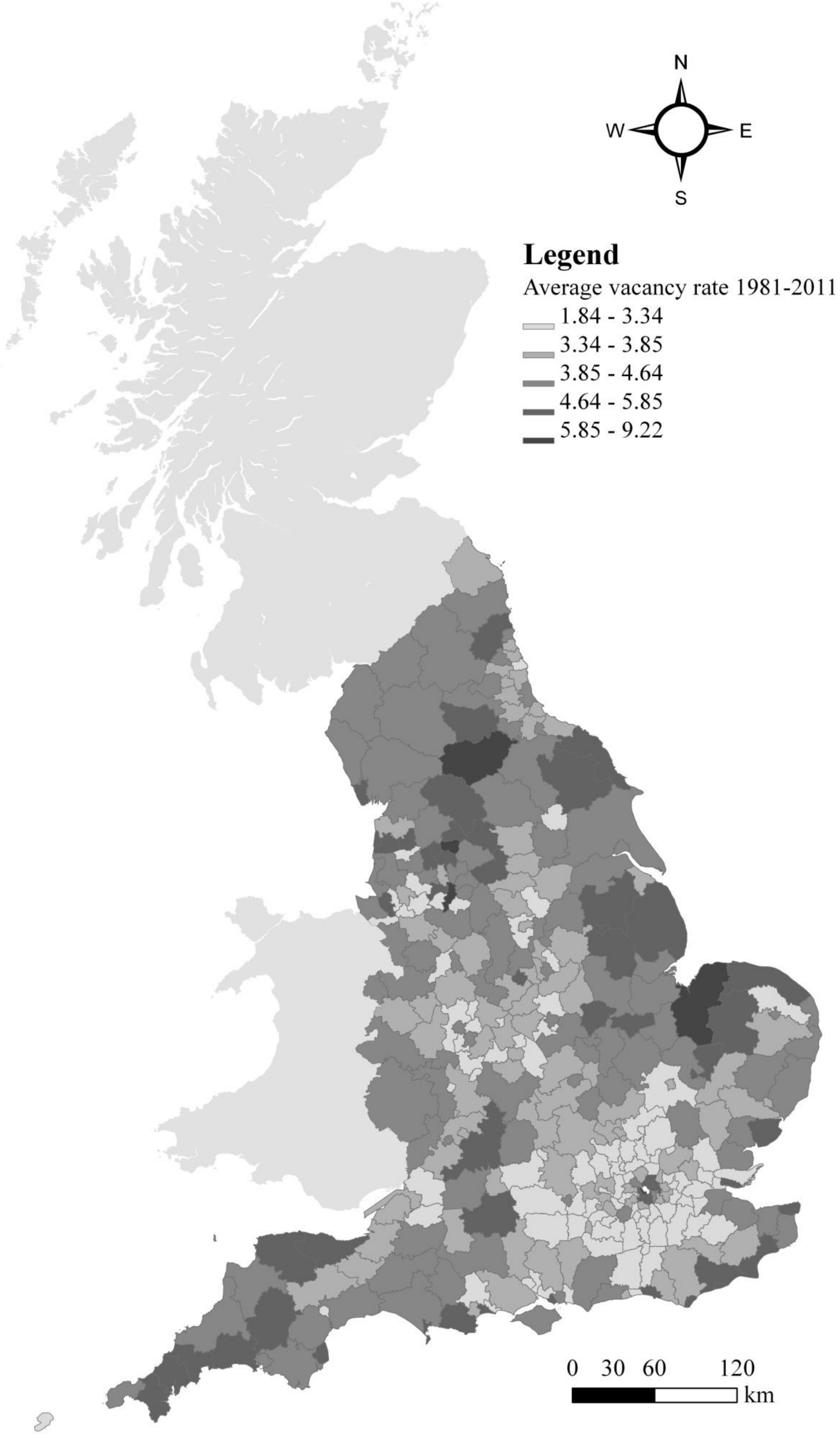


FIGURE 4
Housing supply restrictions across England

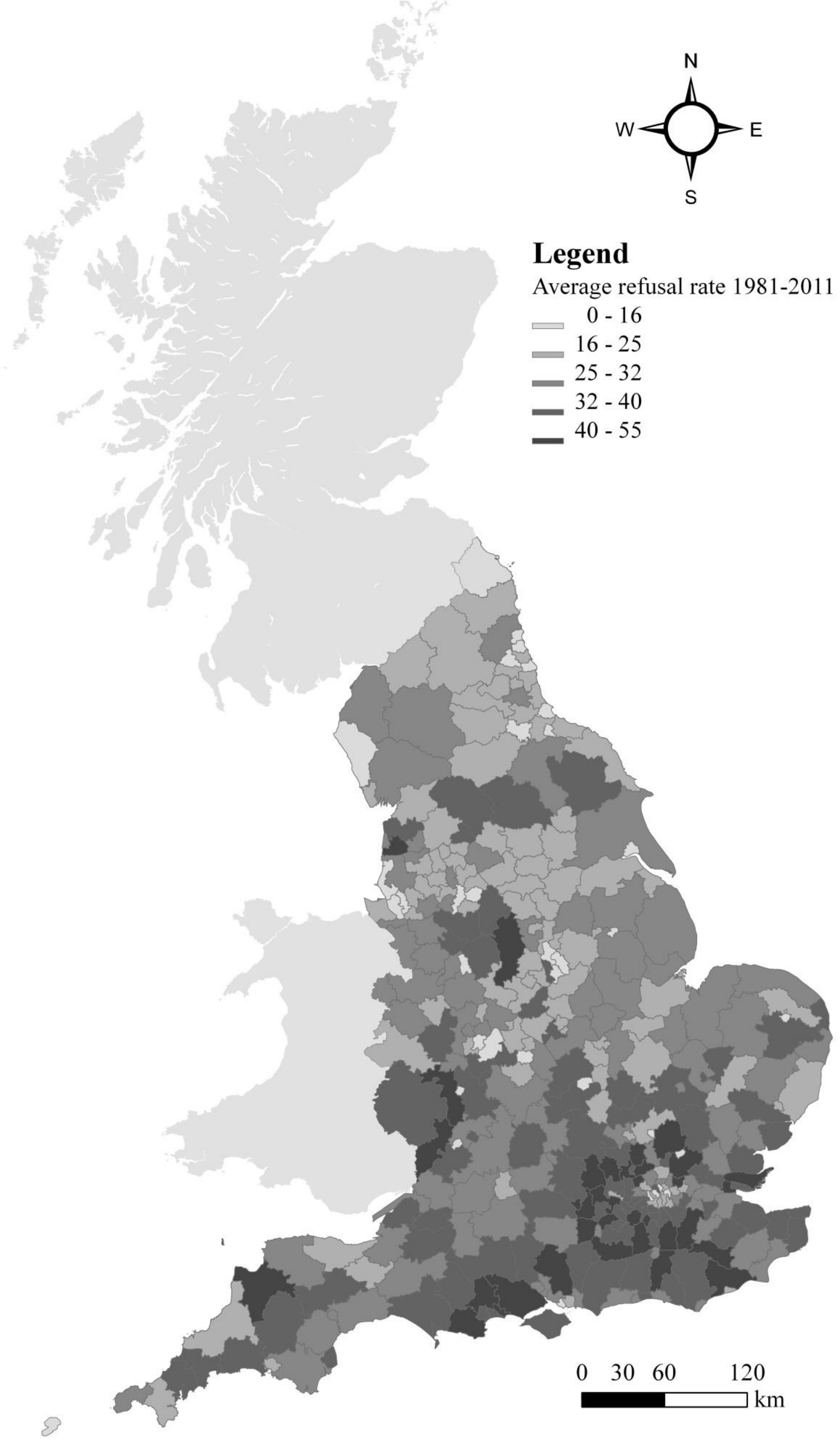
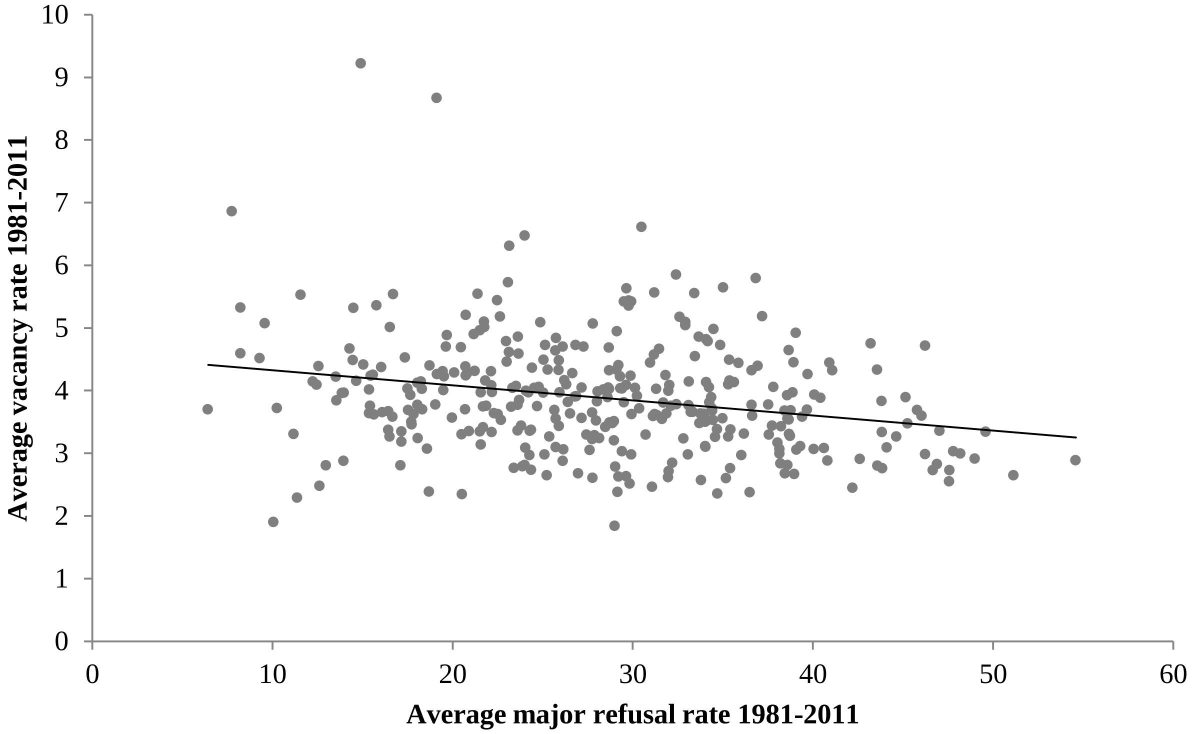


FIGURE 5

Cross-sectional relationship between refusal rate and vacancy rate



APPENDICES

APPENDIX 1: Election data

The local election data are obtained from three different sources. The first source is the British Local Election Database, which is compiled by Rallings and Thrasher (2004). They have combined different data sources on local election outcomes from 1889 to 2003. From 1973, the data contain the universe of local election outcomes. The data is available on a ward level and displays for every election the number of candidates, the number of votes per candidate and the number of vacant seats. Councillors that received the most votes will be elected. It is important to note that the share of votes for each party is therefore not perfectly correlated to the assigned number of seats. Based on the number of votes, we determine which candidate is elected as a councillor. The British Local Election Database only provides information on the election results, and not on the current composition of the local council. The problem is that for many local authorities, there are yearly or two-yearly elections of which 33 percent of the seats are replaced. To estimate the composition of the local council, we use the fact that the full electoral term for *councillors* is usually four years. However, sometimes elections replace the complete council despite the fact that councillors did not complete the electoral term, for example due to changes in boundaries of local authorities. To account for this, we consider full elections as elections where at least 75 percent of the seats are replaced.

The second data source for local election results is the local election handbooks from 1999 to 2008 (see Hilber et al., 2011). These data, also used in provide the number of council seats for each local authority in each year. We measure the correlation of the shares of seats of different parties (Labour, Conservatives, Liberal Democrats, Other) between the British Local Election Database and the latter database for the overlapping years. This is always above 0.95. For the overlapping years, we take the average of shares in seats in both datasets. For 2006-2011, we obtain information on Labour votes (rather than seats) from the Local Elections Archive Project (LEAP). For 55 local authorities, we do not have information available for the most recent election in 2011. We then use information on Labour vote shares for the 2007 elections. We made sure that excluding these 55 local authorities lead to essentially the same results. The final dataset is from the BBC with the outcomes of local council elections for 2009, 2010 and 2011 to complement the LEAP when necessary.

We also use outcomes of general elections to control for demographic changes and general trends in political preferences. We have data of election results for 1983, 1987, 1991, 1997, 2001, 2005 and 2010 obtained from Electoral Calculus. We match each year to the previous election, except for 1981, which is matched to 1983. The results are available at the parliamentary constituency level, which are almost always smaller than local authorities. Using geographical information systems, we calculate the geographical overlap of each constituency with each local authority and assign the votes accordingly.

Table A1 below shows some correlation for the election variables of the Labour party. It is shown that the correlation between the share of Labour council seats and the share of Labour votes is high (0.87). The correlation with general election vote share is somewhat lower, but still reasonable high (0.67). If we look at the correlation between the changes, these are lower (respectively 0.57 and 0.48).

TABLE A1
Correlations between election variables

	Share labour seats	Share labour votes	Share labour votes gen. elec.
Share labour seats, *t-2*	1.0000		
Share labour votes	0.8761	1.0000	
Share labour votes general elections	0.6901	0.6060	1.0000

	Δ Share labour seats	Δ Share labour votes	Δ Share labour votes gen. elec.
Δ Share labour seats, *t-2*	1.0000		
Δ Share labour votes	0.5706	1.0000	
Δ Share labour votes general elections	0.4474	0.3834	1.0000

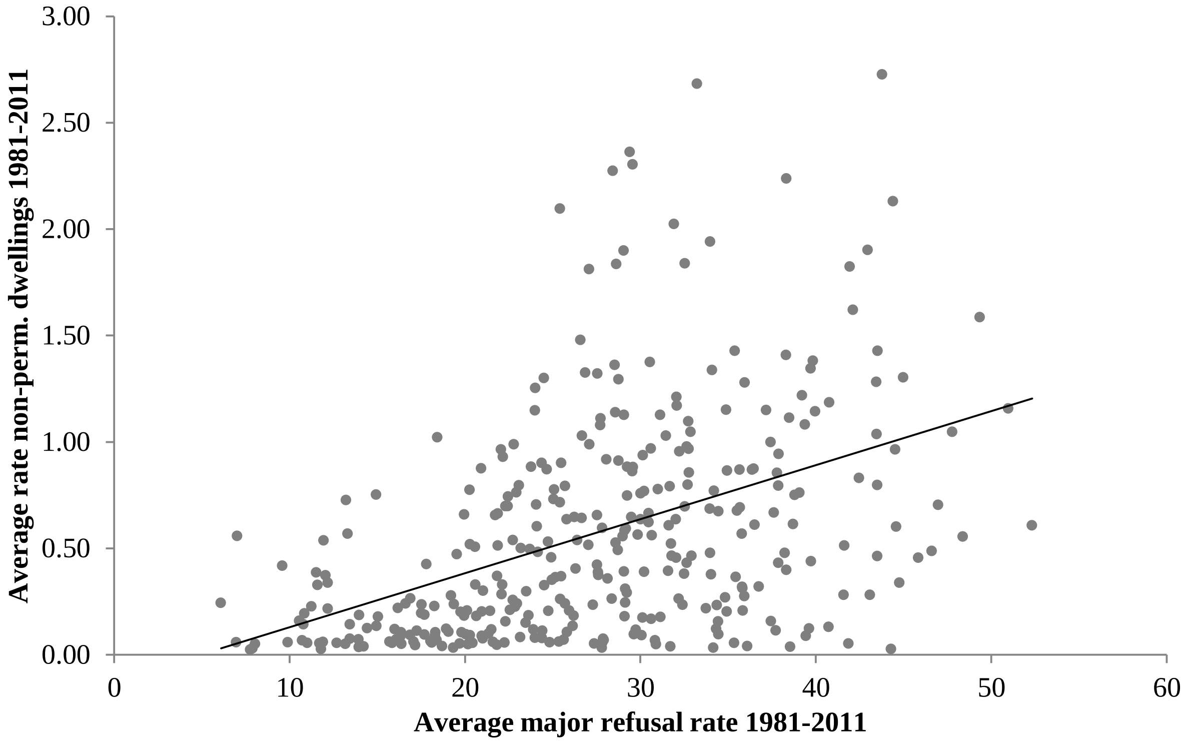
APPENDIX 2: Mismatch and restrictiveness

Measures of mismatch in the local housing stock are not easy to devise. Moreover any relationship between such a measure – however valid it might be – and the restrictiveness with which planning policy is applied locally could be subject to a number of different factors and endogeneity is an issue. In the core analysis of our paper we proposed 'commuting distance from the workplace' as a plausible proxy for the level of mismatch in the housing market. In this appendix we derive an additional (or alternative) indicator: the proportion of the local stock of dwellings which are not permanent (caravans, mobile homes, etc.).

One would unambiguously expect the share of such non-permanent dwellings to increase with search costs in the local housing market since the ease with which search can be undertaken will increase if the house-hunter can be physically present in the local market (Ha and Hilber, 2013). Moreover since the chances of finding a better match in the housing market will improve with length of time spent searching then there will be a payoff to having temporary accommodation available for searchers in markets where matching is more difficult. (Note also that non-permanent homes are not included when calculating the vacancy rate, which is only based on permanent homes.) However since the attraction of temporary homes is that they give cheap access (transactions costs are low compared to buying and then selling houses that do not match the individuals' preferences) to search and so improve its efficiency it is possible that more temporary homes in a more restricted local market may reduce vacancies ceteris paribus. In this sense more temporary homes is likely an indicator of more mismatch rather than a measure.

Figure A1 provides some evidence in support of this proposition. It plots the mean share of non-permanent dwellings (caravans, mobile homes, etc.) over the period 1981-2011 against the major refusal rate in the LA over the same period. There is a strong and positive correlation between these two variables (ρ = 0.425).

FIGURE A1
Non-permanent dwellings and regulatory restrictiveness



We also estimate regressions where we use first-differencing and regress the rate of non-permanent dwellings on the major refusal rate. Table A2 below reports the results of these additional regressions, where we follow the same approach as in Table 2. In column (1) we regress the change in the share of non-permanent dwellings on just the change in major refusal rate. It appears that there is no statistically significant impact of restrictions on the share of non-permanent dwellings. Also if we include local authority fixed effects and demographic controls, the results are statistically insignificant.

However, developers may be less likely to apply at all when there is a high probability that their application will be turned down. This may lead to a bias towards zero of the estimated coefficients. In column (4) we aim to account for the measurement error in the refusal rate: we instrument with the share of Labour seats (as in the rest of the paper). Note that the first-stage results are identical to the ones presented in Table 3. The results in column (4) reveal that the impact of regulatory restrictiveness on the rate of non-permanent dwellings becomes substantially stronger: one standard deviation increase in the major refusal rate leads to an increase in the rate of non-permanent dwellings of 0.09 percentage points (about one-fifth of a standard deviation). Hence, measurement error seems to be important here. However, the result is still not statistically significantly different from zero (p-value = 0.219). In column (5) we further include local authority fixed effects. The point estimate is very similar to the previous specification and statistically significantly different from zero at the 16 percent level (p-value = 0.158). Hence, the estimate is quite imprecisely estimated. This is not too surprising given that the share of non-permanent home is quite noisy. For example, the proportion of temporary homes is very small in the great majority of LAs but in tourist areas it is subject to very considerable variation when for example a caravan park opens (or closes).

In column (6) we include demographic control variables. This produces a similar point estimate but is also statistically significant at the 20 percent level (*p*-value = 0.151). In the final column, we control for a flexible function of the share of labour votes in a local authority. The coefficient then becomes somewhat stronger: one standard deviation increase in the major refusal rate leads to an increase in the rate of non-permanent dwellings of 0.18 percentage points. The result is then close to being statistically significant at the 10% level (*p*-value = 0.114).

Although these regressions do not provide conclusive evidence, the positive relationship between the share of non-permanent dwellings and local restrictiveness does seem to provide some circumstantial evidence. The positive relationship is certainly consistent with the proposition that more restrictive local planning increases the costs of matching would-be house buyers in the local housing market to the available permanent housing, inducing people to live temporarily in caravans and mobile homes – clearly inferior substitutes to houses.

Finally, we include commuting distance and price volatility as additional controls in the regression of vacancy rates on the major refusal rate. Table A3 reports the results, where column (7) in Table 2 is the corresponding specification. In column (1) of Table A3 we show that commuting distance is positively associated with higher vacancy rates. A one kilometre increase in the average commuting distance is associated with an increase in the vacancy rate of about 2 percentage points. We do not interpret this as a strictly causal effect, but the sign is in line with our expectations. The effect of restrictiveness is about 15 percent lower compared to the baseline specification. The effect is now statistically significantly different from zero only at the 13 percent level (*p*-value = 0.131). Commuting distance is a crude proxy for mismatch in the housing market, which may explain why the effect of interest is only somewhat smaller. In column (2) we include the coefficient of variation of prices in the year of observation and the three years subsequent to the year of observation. The coefficient is positive and highly statistically significant. Interestingly, the coefficient on regulatory restrictiveness is now somewhat higher compared to the baseline specification. Column (3) includes both commuting distance and price volatility, leading to similar conclusions. Column (4) also includes the (change in the) share of non-permanent homes. The coefficient on non-permanent homes is positive, in line with expectations but is not statistically significantly different from zero at conventional significance levels. Including the (change in the) share of non-permanent homes reduces the coefficient on the regulatory restrictiveness variable somewhat, in line with the proposition that the positive causal effect from regulatory restrictiveness on housing vacancies is driven by mismatch between contemporaneous household preferences and the stock of housing currently on offer.

TABLE A2

Additional evidence for mismatch in the housing market

(Dependent variable: Δ Rate non-permanent homes)

	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS	(8) 2SLS
Δ Major refusal rate, *t-2*	0.00473	0.00346	0.00525	**0.0902**	**0.107**$^{+}$	**0.104**$^{+}$	**0.181**$^{+}$	**0.207**$^{+}$
	(0.0106)	(0.0135)	(0.0129)	**(0.0734)**	**(0.0761)**	**(0.0721)**	**(0.114)**	**(0.129)**
Δ House prices *(log)*								0.226*
								(0.120)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	No	Yes	Yes	No	Yes	Yes	Yes	Yes
Control variables included (8)	No	No	Yes	No	No	Yes	Yes	Yes
Δ Share labour voters local elections $\Omega(\cdot)$	No	No	No	No	No	No	Yes	Yes
Observations	1,050	1,050	1,050	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.034	0.327	0.362					
Kleibergen-Paap *F*-statistic				18.52	17.49	19.01	8.151	6.681

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. In Models (4)-(7), the instrument for Δ Major refusal rate is Δ Share of labour seats in the local council. $\Omega(\cdot)$ is approximated by a fifth-order polynomial of share labour voters. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10, $^{+}$ p<0.20.

TABLE A3

Results controlling for commuting time and house price volatility

(Dependent variable: Δ Vacancy rate)

PANEL A – Second stage *(Dependent variable: Δ Vacancy rate)*	Baseline 2SLS	(1) 2SLS	(2) 2SLS	(3) 2SLS	(4) 2SLS
Δ Major refusal rate, *t-2*	**0.895***	**0.775**$^{+}$	**1.051****	**0.937***	**0.870***
	(0.530)	**(0.513)**	**(0.517)**	**(0.497)**	**(0.492)**
Δ Commuting distance from workplace *(log)*		1.977***		2.095***	2.021***
		(0.660)		(0.670)	(0.678)
Δ Coefficient of variation of house prices, *t+3*			2.970**	3.225***	3.502***
			(1.172)	(1.107)	(1.093)
Δ Share non-permanent homes					0.478
					(0.317)
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	Yes	Yes	Yes	Yes	Yes
Control variables included (8)	Yes	Yes	Yes	Yes	Yes
Δ Share labour voters local elections Ω(·)	Yes	Yes	Yes	Yes	Yes
Observations	1,050	1,050	1,050	1,050	1,050
Kleibergen-Paap *F*-statistic	8.151	8.307	9.049	9.192	8.078

PANEL B – First stage *(Dependent var.: Δ Major refusal rate, t-2)*		(1) OLS	(2) OLS	(3) OLS	(4) OLS
Δ Share labour seats, *t-2*	-0.264**	-0.283**	-0.275**	-0.293**	-0.290**
	(0.114)	(0.114)	(0.115)	(0.114)	(0.114)
Δ Commuting distance from workplace *(log)*		-0.329		-0.438	-0.441
		(0.315)		(0.326)	(0.327)
Δ Coefficient of variation of house prices, *t+3*			0.165	0.589	0.632
			(0.783)	(0.811)	(0.816)
Δ Share non-permanent homes					0.0825
					(0.186)
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	Yes	Yes	Yes	Yes	Yes
Control variables included (8)	Yes	Yes	Yes	Yes	Yes
Δ Share labour voters local elections Ω(·)	Yes	Yes	Yes	Yes	Yes
Observations	1,050	1,050	1,050	1,050	1,050
Adj. R-squared	0.331	0.327	0.326	0.328	0.328

Notes: **Bold** indicates instrumented. The instrument for Δ Major refusal rate is Δ Share of labour seats. Standard errors are clustered at the LA level and in parentheses. *** p<0.01, ** p<0.05, * p<0.10, $^{+}$ <0.20.

WEB APPENDICES

WEB APPENDIX 1: Theoretical model

W1.1 Matching and sales time

We analyse a setting with search frictions where households may move, but not freely, between housing markets. We extend the seminal house price bargaining model by Wheaton (1990), which can be interpreted as a model with two symmetric housing markets. We follow Koster and Van Ommeren (2015) by relaxing the symmetry assumption. This is needed to analyse the effects of more stringent land use regulations in one of the two housing markets. We first outline the model and discuss the relationship between search effort, matching rates, house prices and regulatory restrictiveness. Using numerical simulations, we then illustrate that regulatory restrictions may both increase or decrease the (relative) vacancy rate, while they will lead to relatively higher prices and lower matching rates.

Assume a country with a given number of home-owning households equal to $2\overline{H}$. The country consists of two local housing markets each with S_i identical housing units, where $i = 1{,}2$. Households own at least one house but may own two houses. All houses, including vacant ones, are owned by a household implying that $(\bar{S} - \overline{H})/\overline{H} < 0.5$. Households prefer to live in one of the two housing markets. Let H_i denote the number of households who prefer to live in housing market i, where $i = 1, 2$. Households who live in their preferred housing market are matched, otherwise they are mismatched. Mismatched households search for housing in the other housing market. Dual-ownership households have a house in both housing markets, are therefore matched, but own a vacant house in the other housing market.

Preferences for a housing market change over time at an exogenous rate β. For example, the location of the job may be exogenously changed from one housing market to the other due to a firm relocation (Mulalich *et al.*, forthcoming). The change over time in H_i denoted by $\dot{H}_i = \beta(H_j - H_i)$ for $i \neq j$. In what follows, j will always denote the other housing market. The total endogenously determined number of households that have a preference to reside in i is given by $H_i = H_i^M + H_i^D + H_i^S$, where H_i^M is the number of matched households with one house, H_i^D is the number of households possessing two houses, and H_i^S denotes the number of mismatched households.

By construction, the vacancy rate in a housing market v_i equals housing supply S_i minus the number of households living there divided by the number of housing units in i. Because mismatched households j live in housing market i, the following holds:

(1) $$v_i = \frac{H_j^D}{S_i} = \frac{S_i - H_i^D - H_i^M - H_j^S}{S_i}.$$

Each vacant house is owned by a household that lives in the other housing market, so $H_i^D = v_j S_j$ where $i \neq j$. Hence, the number of vacant housing units in housing market j is equal to the number of dual-ownership households living in housing market i. Mismatched households search for houses in the other housing market and given a contact with a vacancy in that housing market, they find a (suitable) house with probability one. Matching of mismatched households and vacant houses, and therefore the sales of vacant residences, occur with a Poisson process. The sales rate q_i equals the product of the number of mismatched households of type i and m_i, the matching rate of a mismatched household i, divided by the number of vacancies in housing market i:

(2) $$q_i = \frac{m_i H_i^S}{V_i} = \frac{m_i}{\theta_i},$$

where $\theta_i = V_i / H_i^S$, so the ratio of vacancies to mismatched households. The sales time is then defined as $\ell_i = 1/q_i$. Given the above assumptions, following Wheaton (1990), the first-order differential equations that indicate how households change type and move between housing markets are given by:

(3) $$\dot{H}_i^S = -(m_i + \beta)H_i^S + \beta H_j^M,$$

(4) $$\dot{H}_i^D = -(q_j + \beta)H_i^D + m_i H_i^S + \beta H_j^D,$$

(5) $$\dot{H}_i^M = -\dot{H}_i^S - \dot{H}_i^D, \qquad i \neq j$$

The above equations provide a stable model of changes in household type and moving.

W1.2 Steady state and regulatory restrictions

We will now assume that the market is in steady-state. Hence, $\dot{H}_i = \dot{H}_j = 0$, and $H_i = H_j = \bar{H}$.[14] We keep the total number of housing units in the housing market fixed and equal to $S^T = 2\bar{H}/(1-n)$, where n may be interpreted as the natural vacancy rate.[15] We then define:

(6) $$S_i = \frac{1-\rho}{1-n}\bar{H} \qquad \text{and} \qquad S_j = S^T - S_i,$$

where ρ indicates the degree of restrictiveness. When $\rho = 0$, the number of housing units in both housing markets is the same. Conditional on the matching rates m_i and m_j (to be determined in the next subsection), we can obtain explicit expressions for the number of households of each type. It can be shown that:

(7) $$H_i^S = \frac{\beta(1-2n+\rho)}{(1-n)(2\beta+m_i)}\bar{H} \qquad \text{and} \qquad H_j^S = \frac{\beta(1-2n-\rho)}{(1-n)(2\beta+m_j)}\bar{H}.$$

Given the latter equality, we obtain the intuitive result that the number of mismatched households in i depends negatively on the matching rate m_i (but does *not* depend on the matching rate of the other housing market m_j).[16] It is also observed that the matching rate in i depends positively on regulatory restrictiveness, while the number of mismatched households in j decreases when ρ is higher. We may also interpret this as an indirect way of commuting distance. If β signals the workplace, a higher H_i^S will imply a longer workplace commute distance from i because mismatched households of type i still live in j.

Similarly, the number of matched households in *i* can be written as:

[14] In addition, the number of households who move residence from one housing market into the other is equal to the number of households who move the other way round, so $m_i H_i^S = m_j H_j^S + \beta(H_i^D - H_j^D)$, where the last term denotes the net number of dual-ownership households who move from i to j, because they have changed preference while holding two houses.

[15] One may endogenise housing supply and therefore n as a function of house prices. We discuss this issue later.

[16] In line with intuition, when m_i approaches infinity, so search frictions in i are absent, the number of mismatched households of type i (who live in *j*) approaches zero.

$$(8) \qquad H_i^M = \frac{(1-2n-\rho)(\beta+m_j)}{(1-n)(2\beta+m_j)}\overline{H} \quad \text{and} \quad H_j^M = \frac{(1-2n+\rho)(\beta+m_i)}{(1-n)(2\beta+m_i)}\overline{H}$$

which is an increasing function of the matching rate in the other housing market, but does *not* depend on the own matching rate. From (8), it can be immediately seen that restrictions reduce the number of matched households in i.

The vacancy rate v_i is equal to the number of dual owners in housing market j divided by the housing supply in i. Using (1), (6), (7) and (8):

$$(9) \qquad v_i = \frac{\beta\left((\rho-1)m_j - 4n\beta\right) + m_i\left((1-4n+\rho)\beta - (n-\rho)m_j\right)}{(\rho-1)(2\beta+m_i)(2\beta+m_j)},$$

It is more informative to evaluate the derivative of the vacancy rate with respect to restrictions:

PROPOSITION 1 [Vacancy rate effect of restrictions]: Restrictions may either positively or negatively affect vacancy rates, $\mathrm{d}v_i/\mathrm{d}\rho \lesseqgtr 0$.

Proof. The derivative of (9) is given by:

$$(10) \qquad \frac{\partial v_i}{\partial \rho} = \frac{\frac{(n-1)m_i}{2\beta+m_i} + \frac{2n\beta}{2\beta+m_j}}{(\rho-1)^2},$$

The denominator is always positive. In the presence of regulations, the vacancy rate will increase when the natural vacancy rate is high relative to the matching rates m_j. Also when $m_i \to 0$ (but m_j is not close to zero), (10) becomes positive, so in inefficient markets with low matching rates, restrictions may lead to increases in vacancy rates. ■

W1.3 Matching and house prices

In the previous subsection, matching rates were given. We will now assume that matching rates are endogenously determined. Mismatched households choose the level of search effort e_i to find a new house. We assume a constant return to scale Cobb-Douglas matching function, following the theoretical and empirical literature on labour market matching:

$$(11) \qquad m_i = \mu e_i^{\psi} \theta_i^{1-\psi},$$

where μ is a constant. Note again that $\theta_i = V_i / H_i^S$.

The level of search effort is based on the households' present discounted value of being in a certain state, which depends, among others, on house prices. In steady state, the value of a certain state can be written as the utility flow of being in that state plus the value of changing state:

$$(12) \qquad ru_i^M = A_i + \beta(u_j^S - u_i^M),$$

$$(13) \qquad ru_i^D = A_i + q_j(u_i^M - u_i^D + p_j) + \beta(u_j^D - u_i^D),$$

(14) $$ru_i^S = -c_i + \beta\left(u_j^M - u_i^S\right) + m_i\left(u_i^D - u_i^S - p_i\right), \quad i \neq j$$

where u_i^M, u_i^D, u_i^S are the present values of each state (matched, dual ownership, mismatched), A_i refers to the utility flow of the housing market-specific amenity, r is the discount rate, c_i is the cost of search and p_i is the house price. Matched and dual-owners households of type i enjoy the housing market-specific amenity A_i, see (12) and (13), whereas mismatched households of type i (who live in j) do *not* enjoy A_j.[17] Households are assumed to choose the level of search effort e_i to maximise u_i^S.

In equation (14), it is assumed that mismatched households incur search costs c_i as a function of search effort. We will assume that $c_i = e_i^2/2$, so c_i is an increasing and convex function of search effort.[18]

Here, we only consider symmetric equilibria where all households choose the same search effort level. The *individual* matching rate of a mismatched household of type i is the product of individual search effort and the average number of matches in the point where $\tilde{e}_i = e_i$, where $\tilde{e}_i$ is the individual level of search effort.[19] Then:

(15) $$\frac{\partial m_i}{\partial \tilde{e}_i} = \frac{m_i}{e_i},$$

The marginal cost from an additional unit of search effort is equal to the marginal benefit of that unit. Given equation (14), optimal choice of search effort implies:

(16) $$\frac{\partial c_i}{\partial \tilde{e}_i} = \frac{\partial m_i}{\partial \tilde{e}_i}\left(u_i^D - u_i^S - p_i\right),$$

where $u_i^D - u_i^S - p_i$ indicates the benefit of a match for mismatched households.

In a general equilibrium setting, prices are endogenous. We assume that buyers and sellers bargain about prices using Nash-bargaining, where buyers get a fixed share σ of the total benefits of the match. So, $u_i^D - u_i^S - p_i = \sigma\left(u_i^D - u_i^S + u_j^M - u_j^D\right)$ and hence $p_i = \sigma\left(u_j^D - u_j^M\right) + (1-\sigma)\left(u_i^D - u_i^S\right)$.

W1.4 Numerical simulations

In this subsection we show that under reasonable parameter values, vacancy rates may be positively associated with regulatory restrictiveness. We assume that $2\overline{H} = 1900$, $\beta = n = 0.15$, $\rho = 0.1$. We furthermore assume that sellers and buyers have equal bargaining power ($\sigma = 0.5$) and the interest rate is 2.5 percent ($r = 0.025$). The amenity A level is set to 10 in both housing markets. We analyse the outcomes when the market is in steady state. The model is solved using an iterative two-step procedure.[20] In Figure W1 we plot the matching rates in i and j as a function of the matching constant μ. Given the first order condition of (16), when μ increases search effort will be higher and matching rates

[17] If one assumes that mismatched households receive a positive share of A_j, results will not change.

[18] We may also assume that dual ownership households incur search costs. However, because we will allow for bargaining on the price later on, this will not change the outcomes of the model, while it unnecessarily complicates matters.

[19] This means that the second-order derivative of the matching rate with respect to *individual* search effort is equal to zero.

[20] First, conditional on an initial (arbitrary) choice of values for number of household variables H_i^S, H_j^S, H_i^M and H_j^M, we solve for house prices, optimal search effort in both areas. Given optimal search effort, we determine m_i, m_j, q_i and q_j. Given these values, we calculate the number of households in each state using equations (3) to (5). We continue this process until the model converges.

are higher because it is more costly to leave a property vacant essentially. It is shown that the matching rate in i is always lower than in j given restrictions in i (the refusal rate is set to 0.1). Figure W2 shows that vacancy rates in i are only higher than in j when the matching constant is smaller. That is, when the opportunity costs of leaving a property empty are not too high. Figure W3 confirms the empirical stylised fact that prices in restrictive local housing markets are above the house prices of less restrictive housing markets. We note that one may also endogenise housing supply as a function of prices. This would mean that, conditional on supply restrictions, supply would increase in the more restrictive housing market i. This means that the vacancy rate in i would increase even more, amplifying the positive effect of restrictions on vacancy rates.

FIGURE W1
Matching rates and restrictions

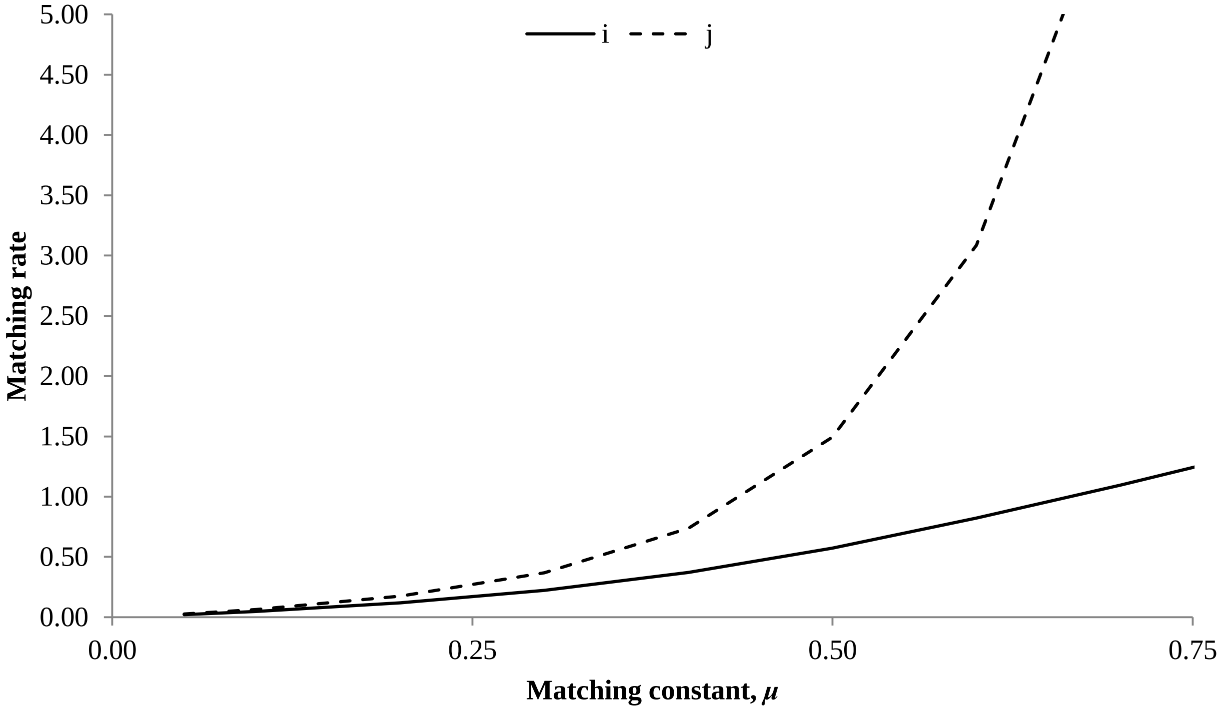


FIGURE W2

Vacancy rates and restrictions

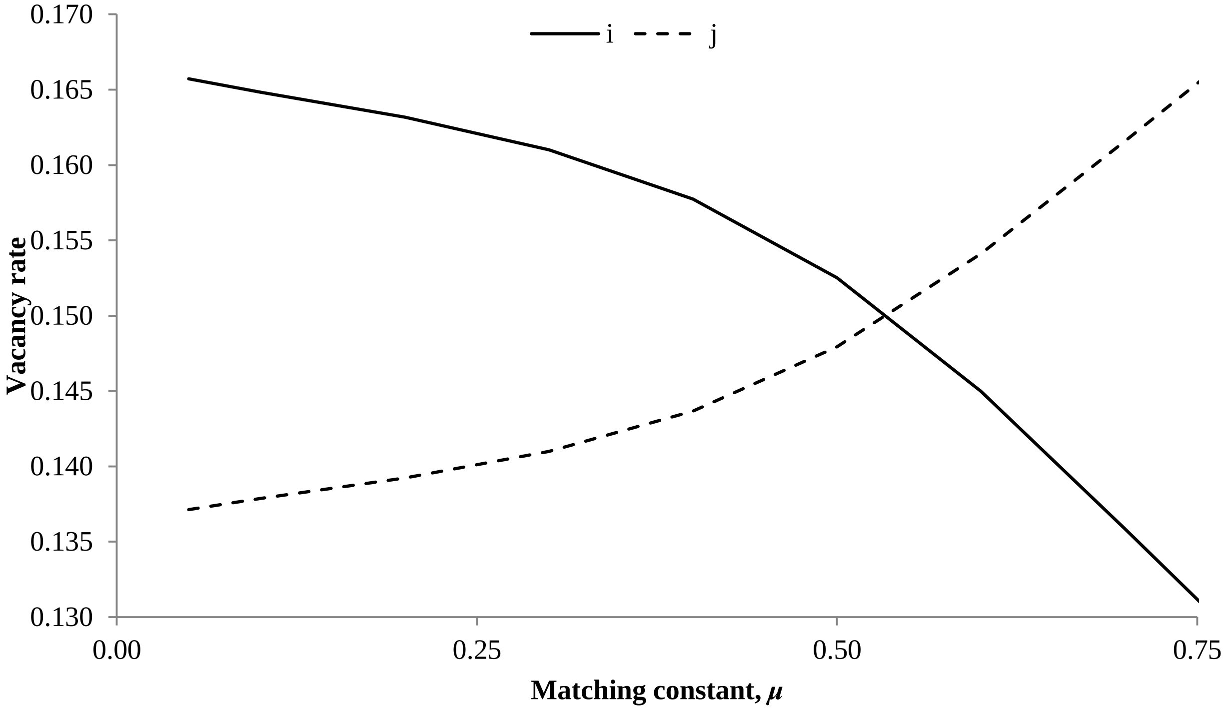
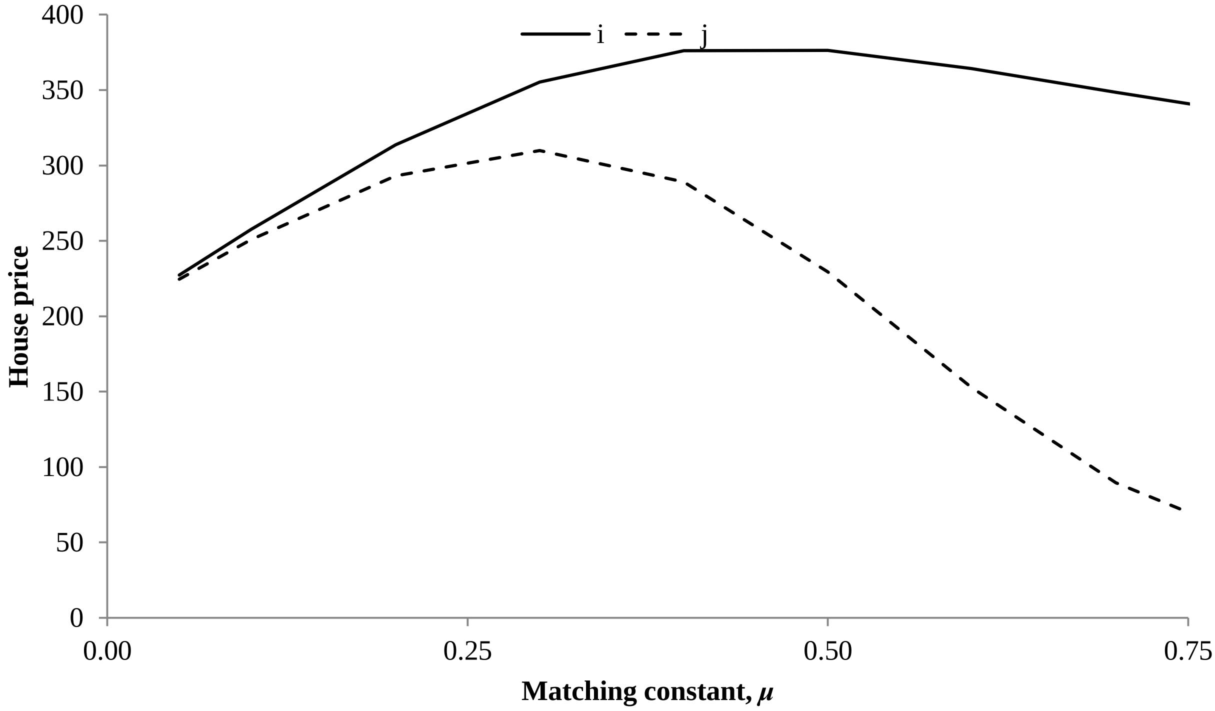


FIGURE W3

House prices and restrictions



WEB APPENDIX 2: Additional robustness checks

In this web appendix we consider additional sensitivity analyses. The results are reported in Table W1. First, we exclude the Greater London area to test whether the results are driven by the restrictive metropolitan area of London. Column (1) in Table W1 shows that the coefficient related to restrictions is even somewhat stronger compared to the baseline specification in column (7), Table 2. Hence, our results are not driven by the Greater London area. In column (2), Table W1, we focus on all restrictions as a sensitivity check. Because minor applications are much less important than major applications; the latter referring to the construction of at least 10 dwellings, while the first may refer to an application to construct an attic. We therefore, somewhat arbitrarily, first calculate the major and minor refusal rate and then take the average to arrive at the total refusal rate. The results indicate that the coefficient related to the total refusal rate is somewhat higher, albeit similar to the baseline specification. Because the minor refusal rate is much noisier than the major refusal rate, the Kleibergen-Paap *F*-statistic is much lower than in the baseline specification. This translates into somewhat less precise second-stage estimates, although the effect is still statistically significantly different from zero at the ten percent level.

Finally, we pursue a fixed effects approach, rather than first-differencing. In column (5) we regress the vacancy rate on the major refusal rate, while controlling for demographic variables, all in levels. We also include LA fixed effects. The results indicate then that a one standard deviation increase in the refusal rate leads to an increase in the vacancy rate of 0.76 percentage points, which is similar to the baseline specification. In column (6) we include 354 local authority-specific linear trends. Results are essentially unchanged. Column (7) includes non-linear trends by estimating second-order polynomials. The effect almost doubles to 1.82, but that may be due to weaker identification (the Kleibergen-Paap *F*-statistic is relatively low with 5.08). Nevertheless, this suggests that controlling more carefully for unobserved time-varying factors of locations, the effect of restrictions on vacancy rates does not disappear.

TABLE W1

Additional sensitivity checks – second stage results

	(1)	(2)	(3)	(4)	(5)
	Exclude Greater London	*Total refusal rate*	*Fixed effects approach – no trends*	*Fixed effects approach – linear trends*	*Fixed effects approach – non-linear trends*
	2SLS	2SLS	2SLS	2SLS	2SLS
Δ Major refusal rate, *t-2*	**1.218****				
	(0.511)				
Δ Total refusal rate, *t-2*		**1.124***			
		(0.673)			
Major refusal rate, *t-2*			**0.757****	**0.714****	**1.822*****
			(0.346)	**(0.294)**	**(0.690)**
Control variables included (8)	Yes	Yes	Yes	Yes	Yes
Δ Share labour voters local elections Ω(·)	Yes	Yes	Yes	Yes	Yes
Share labour voters local elections Ω(·)	Yes	Yes	Yes	Yes	Yes
Local authority fixed effects (350)	Yes	Yes	Yes	Yes	Yes
Local authority-specific linear trends (350)	No	No	No	Yes	Yes
Local authority-specific non-linear trends (350)	No	No	No	No	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	954	1,050	1,400	1,400	1,400
Kleibergen-Paap *F*-statistic	9.608	2.962	17.52	16.63	5.081

Notes: All independent variables (except for earnings) are standardised with mean zero and unit standard deviation. **Bold** indicates instrumented. The instrument for (Δ) Major refusal rate is (Δ) Share of labour seats in the local council. Ω(·) is approximated by a fifth-order polynomial of share labour voters. Standard errors are clustered at the LA level and in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.10$, $^{+}$ $p<0.20$.

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