The London School of Economics and Political Science

Essays in the economics of land, housing, and urban policy

Xiaolun Yu

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Declaration

I certify that the thesis I have presented for examination for the PhD degree of the London School of Economics and Political Science is solely my own work other than where I have clearly indicated that it is the work of others (in which case the extent of any work carried out jointly by me and any other person is clearly identified in it).

I warrant that this authorisation does not, to the best of my belief, infringe the rights of any third party.

I declare that my thesis consists of 50,837 words including footnotes but excluding bibliography and appendices.

Statement of conjoint work

I confirm that Chapter 1 was jointly co-authored with Dr Felipe Carozzi and Professor Christian Hilber and I contributed 50% of this work. Chapter 3 was jointly co-authored with Professor Michael Ball, Professor Paul Cheshire, and Professor Christian Hilber and I contributed 50% of this work.

Xiaolun Yu

Abstract

This thesis combines econometric methods with spatial techniques to study the economics of land, housing, and urban policy in the context of both developed and developing countries. It is organized into four independent chapters. The first chapter (co-authored) takes advantage of two spatial discontinuities in Britain's Help to Buy (HtB) scheme to explore the effectiveness and distributional effects of mortgage credit expansion policies. We find that HtB significantly increased house prices and had no discernible effect on construction volumes in Greater London. We conclude that HtB may be ineffective in already unaffordable and supply constrained areas. The second chapter studies the determinants of floor area ratio (FAR) limit, a major form of land use regulation that specifies construction density, in China. I develop a spatial equilibrium framework to explore the designation process of FAR limit and the trade-off faced by local governments. Exploiting a rich dataset of land transactions and the exogenous variation generated by a central government policy, I find that local budgetary revenue has a negative impact on FAR design. The third chapter (co-authored) explores the determinants of the speed of residential development after the onset of construction. Using a sample of over 110,000 residential developments in England from 1996 - 2015 and employing an instrumental variable- and fixed effects-strategy, we find that strong local demand increases the rate of site build out, but less so for projects located in areas with more restrictive supply constraints and less competition among developers. The last chapter measures 'regulatory tax' in 117 major Chinese cities by using a spatially matched dataset of land plots and residential projects. The measure shows substantial variations in regulatory restrictiveness across Chinese cities. I also find that housing prices respond more strongly to local demand shocks in cities with more severe regulatory and geographical constraints.

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Preface

This thesis contains four independent chapters on the economics of land, housing, and urban policy in the context of both developed and developing countries.

The first chapter, 'On the economic impacts of mortgage credit expansion policies: evidence from help to buy', takes advantage of two spatial discontinuities in Britain's Help to Buy (HtB) scheme to explore the effectiveness and distributional effects of mortgage credit expansion policies. Employing a Difference-in-Discontinuities design, we find that HtB significantly increased house prices and had no discernible effect on construction volumes in severely supply constrained and unaffordable Greater London. Conversely, HtB did increase construction numbers without affecting prices near the English/Welsh border, an affordable area with lax supply conditions. HtB also boosted the financial performance of developers participating in the scheme. We conclude that HtB may be ineffective in already unaffordable and supply constrained areas.

The second chapter, 'Low-rise buildings in big cities: theory and evidence from China', studies the determinants of floor area ratio (FAR) limit — a major form of land use regulation that specifies construction density — in China. I first develop a spatial equilibrium framework that assumes that local governments set FAR limits such as to maximize endogenous local GDP and population size. Setting a higher limit enables them to provide more public goods and more housing but also increases negative externalities such as more congestion and pollution. I show that in equilibrium, cities with higher TFP can collect more budgetary revenue and opt to set lower FAR limits going forward to reduce negative externalities caused by density. I then employ a rich dataset of over 200,000 residential land transactions in China and a county-level panel to test for the theory's prediction. I exploit the exogenous variation generated by a central government administrative adjustment policy and a propensity score matching approach and find that a one standard deviation increase in local government budgetary revenue decreases FAR limits by 0.6. I conclude that the Chinese 'Land Finance Model' influences local land use design and contributes to the country's housing affordability issues and regional inequality.

In the third chapter, 'Why delay? Understanding the construction lag, AKA the build out rate', we explore the determinants of the speed of residential development after onset of dwelling construction. Using a sample of over 110,000 residential developments in England from 1996 - 2015 and employing an instrumental variable- and fixed effects-strategy, we find that strong local demand increases the rate of site build out, but less so for projects located in areas (i) where local planning is more restrictive, (ii) that are already built-up, and (iii) where competition in the local development sector is lower. This suggests that restrictive local planning and market concentration adversely affect the elasticity of housing supply even after planning permission is granted.

The last chapter, 'Supply constraints and housing markets: evidence from a spatially matched dataset', applies and extends the method proposed by Glaeser *et al*. (2005) to estimate 'regulatory tax rate' in 117 major Chinese cities. Using a unique spatially matched dataset of land plots and residential projects, I find substantial spatial variations in regulatory restrictiveness across Chinese cities. I then study the impacts of both regulatory constraint and geographical constraint on housing prices by estimating a city-level panel dataset between 2005 and 2018. Exploiting the exogenous variation generated from local dialects, government tax enforcement, and historical amenities, I find that real housing prices respond more strongly to changes in local real salaries in cities with more restrictive supply constraints. This finding is robust after I apply a Bartik-type predicted local employment as the demand shifter. I also find that housing construction and land supply respond less strongly to changes in local salaries in places with tighter supply constraints.

Although each chapter approaches a different research question, the unifying aim of this thesis is to understand the functioning of land and housing markets by combining rigorous identification strategies and spatial techniques with solid theoretical grounds. This thesis strives to contribute to the literature on urban and real estate economics by exploring the effectiveness of mortgage credit expansion policies, investigating the determinants of land use regulation and construction lag, and documenting the impacts of supply constraints on housing markets.

Chapter 1

On the Economic Impacts of Mortgage Credit Expansion Policies: Evidence from Help to Buy

(with Felipe Carozzi and Christian Hilber)

1. Introduction

Government policies directed at stimulating demand or supply in mortgage markets are widespread throughout the world. Examples of credit market interventions include mortgage interest deductions in the United States, India or Sweden, mortgage guarantees in the United States or the Netherlands, and government loans for home purchases in France or the United Kingdom. Most of these interventions have the explicit goal of making homeownership more affordable and thus accessible. In a context in which housing affordability problems are increasingly pervasive, especially in large desirable cities, new policies are discussed – and implemented – frequently.

Many recent housing and credit expansion policies also tie homeownership subsidies to the purchase of newly built homes. Examples include policies implemented in countries as diverse as Australia, Ireland, Mexico, and Lithuania, in addition to the policy investigated in our paper. Governments thereby seek to stimulate new housing construction, which in turn should further aid making owner-occupied housing more affordable.

In this paper, we employ a unique setting – spatial discontinuities in an equity loan scheme that is tied to the purchase of newly built homes[1] – to shed light on the effectiveness and distributional effects of such policies.

In April 2013, the British government launched a new flagship housing credit policy: Help to Buy (HtB). The program was initially implemented in England, but Welsh and Scottish versions were put in place shortly thereafter. We set out to explore the causal

[1] Policies that tie homeownership subsidies to the purchase of new build properties are increasingly widespread (OECD 2019). However, the research on this type of policies is scant. One exception is Agarwal *et al.* (2020) who study the impact of a stamp duty relief on purchases of newly built homes in Sydney, Australia, finding significant bunching just at and below the threshold price up to which the relief was granted.

impact of HtB on housing construction, house prices, the size of newly constructed units, and the financial performance of residential developers who participated in the scheme. To do so, we focus on the HtB 'Equity Loan Scheme', which provides an equity loan for up to 20% of the housing unit's value – or 40% within the Greater London Authority (GLA) – to buyers of new build properties. The Equity Loan Scheme is by far the most salient and popular of the four HtB schemes and the one requiring the biggest budget. It is often referred to simply as "Help to Buy" and henceforth, unless we note otherwise, when we refer to HtB we mean the Equity Loan Scheme.[2]

HtB expands housing credit and thus increases demand for housing. To explore how such a positive demand shock in the housing market affects construction and prices, we develop a simple theoretical framework with heterogeneous households and credit constraints. Our model predicts that the impact of the policy depends crucially on the responsiveness of supply to prices. In a setting with responsive supply, HtB can be expected to mainly stimulate construction numbers as intended by the policy. However, when supply is unresponsive (i.e., regulatory constraints or physical barriers to residential development impede a supply-response), the effect of the policy may be mainly to increase house prices, with the unintended consequence of making housing less rather than more affordable.

In our empirical analysis, we implement a Difference-in-Discontinuity design to compare changes in house prices and construction activities across jurisdictional boundaries. We separately analyse properties sold on either side of the GLA boundary and on either side of the English/Welsh border. In both cases we only consider housing purchases close to the respective boundaries. In Wales the scheme only applied to a subset of the properties that were eligible in England. Likewise, the London scheme that was implemented in 2016 offered larger government equity loans (as a share of house values) for dwellings inside the GLA compared to those available for purchase outside this region. Our main estimates exploit these spatial discontinuities to study

[2] At the time of implementation, HtB consisted of four schemes: the Equity Loan Scheme, Mortgage Guarantees, Shared Ownership, and Individual Savings Accounts (ISA). All four schemes aim to help credit constrained households to buy a property. The Mortgage Guarantees scheme ceased at the end of 2016. The HtB-ISA closed for new entrants in November 2019 and any bonus must be claimed by 2030. In April 2017, the British government introduced a new Lifetime ISA scheme. In contrast to HtB ISA, it is only open to individuals aged 18-39 and the money saved can also be used to fund a pension.

the effect of HtB on house prices and construction activity. We also use this design to study the impact of the scheme on the size of newly constructed units.

We focus on the GLA boundary and the English/Welsh border for two reasons. First, our research design requires spatial discontinuities in the scheme's conditions, which can be found at these boundaries. Second, the two areas differ starkly in their regulatory land use restrictiveness and in barriers to physical development: While the GLA is the most supply constrained and the least affordable area in the UK – and arguably one of the most supply constrained areas in the world – housing supply is comparably responsive to demand shocks near the English/Welsh border.

Consistent with our theoretical predictions, we find that differences in the intensity of the HtB-treatment have heterogeneous effects depending on local supply conditions. In the GLA, the introduction of the more generous London version of the Equity Loan Scheme led to a significant increase in prices for new build units of nearly 8%. However, it had no appreciable effect on construction activity. Conversely, in the areas around the English/Welsh border, we find a significant effect on construction activity but none on prices. The introduction of the more generous HtB-price threshold on the English side of the border increased the likelihood of a new build sale by about 8% compared to the Welsh side of the border. We also provide evidence indicating that the scheme may have caused an improvement of the financial performance of developers: larger revenues as well as higher gross and net profits.

Collectively, these results suggest that the effects of HtB largely depend on local supply conditions. We find that the scheme fails to trigger more construction activity, but instead causes house prices to increase inside the GLA, precisely the region that is most affected by the 'affordability crisis'. This has distributional implications. The main beneficiaries of HtB in already unaffordable areas may be developers and landowners rather than struggling first-time buyers. While access to homeownership is improved in principle (credit constraints are relaxed), the present value of the financial burden associated with the purchase of a home further increases. In fact, house prices of newly built units inside the GLA increased by about twice as much as the interest rate subsidy induced by the favourable rates on HtB lending.

Our paper relates to previous studies looking at the effects of credit conditions and credit market policies on housing markets. Previous research has mainly focused on the effect of credit supply on house prices (see Ortalo-Magné and Rady 1999, Mian and Sufi 2009, Duca *et al.* 2011, Favara and Imbs 2015, Justiniano *et al.* 2019). These and other studies provide theoretical and empirical credence to the notion that expansions in credit supply may lead to higher prices.

On the policy evaluation front, a few studies have explored the impact of demand subsidies on housing market outcomes. Hilber and Turner (2014) examine the impact of the U.S. mortgage interest deduction (MID). They find that the MID boosts homeownership attainment only of higher income households in markets with lax land use regulation. In tightly regulated markets with inelastic long-run supply of housing, the MID lowers homeownership attainment, presumably because higher house prices also raise down-payment constraints of would-be-buyers. Sommer and Sullivan (2018) estimate a dynamic structural model of the housing market to study the effect of removing the MID and predict this would result in a substantial reduction in house prices. Finally, a related literature has explored the effect of credit expansion policies in the US – such as FHFA guarantees and GSE lending – on homeownership attainment, finding mixed results.[3]

Our paper is the first to provide a state-of-the-art evaluation of the causal impacts of credit expansion policies on house prices, construction volumes and the financial performance of developers. We are also the first to emphasize that the distributional effects of credit expansion policies are spatially heterogeneous in that they depend on local supply conditions.

We are not the first to study the HtB policy. Finlay *et al.* (2016), combining qualitative and non-causal quantitative methods, estimate that since its introduction HtB has generated 43% additional new homes. Szumilo and Vanino (2018) provide evidence that HtB increased the lending volume in Greater London. In a similar vein, exploiting geographic variation in exposure to HtB, Tracey and van Horen (2021) find that HtB increased home sales and consumption of non-housing related items. Benetton *et al.* (2018) utilize the HtB equity loan scheme to investigate the pricing of mortgage credit, demonstrating that a lower down-payment is associated with a higher interest rate at origination, and a higher ex-post default rate. Lastly, Benetton *et al.* (2019) explore

[3] See for example Bostic and Gabriel (2006), Gabriel and Rosenthal (2010) and Fetter (2013). Olsen and Zabel (2015) review the US literature. Hilber and Schöni (2016) provide a comparison of US policies with policies in the UK and Switzerland. An evaluation of the French *Pret a Taux Zero* policy – which provides a down-payment subsidy to low and middle-income first-time buyers – can be found in Gobillon and le Blanc (2008).

the effect of HtB on households' house purchase and financing decisions. They find that households take advantage of an increase in the HtB maximum equity limit to buy more expensive properties.

Finally, this paper links to previous research on housing and land supply, including work on the effects of supply constraints on the responsiveness of housing markets to economic shocks (Hilber and Vermeulen 2016), the origin of supply restrictions (Saiz 2010, Hilber and Robert-Nicoud 2013) and their consequences (see Gyourko and Molloy 2015 and the references therein). We contribute to this literature by studying in depth the effect on housing supply of arguably the most important new British housing policy since the implementation of Right to Buy in 1980.

The rest of this paper is structured as follows. Section 2 describes the details of the HtB Equity Loan Scheme and provides a simple theoretical framework to guide the empirical analysis. Section 3 presents the data sources and discusses the housing supply conditions in the two regions we use in our empirical analysis. Section 4 discusses our empirical strategy in detail and presents our main results for the effect of the policy on housing prices, construction volumes and developer performance. Section 5 provides robustness checks for these findings, while Section 6 presents additional results on price bunching and the effect of HtB on the size of newly built units. In Section 7 we gauge the magnitude of the HtB-interest subsidy and compare it to the HtB-induced increase in the price of newly built homes. Section 8 concludes.

2. Background and Theoretical Framework

2.1. Background: The Help to Buy Equity Loan Scheme

Since the launch of HtB in April 2013 until December 2019, over 263,000 properties in England were bought with a government equity loan provided by the scheme. The total value of these loans is £15.3 billion, with the value of properties purchased under the scheme totalling £70 billion (Ministry of Housing, Communities and Local Government 2020).[4]

[4] The Ministry of Housing, Communities and Local Government (2020) provides a comprehensive overview and numerous summary statistics relating to the HtB Equity Loan Scheme in England. In the case of Wales, since the launch of the Welsh HtB in January 2014 until December 2019, 10,215 properties were bought with a government equity loan (StatsWales 2020).

The English version of the HtB Equity Loan Scheme offers government loans of up to 20% of a unit value to households seeking to buy a new residence. It is available to both first-time buyers and home-movers but it is restricted to the purchase of new build units with prices under £600,000.

Given that the prevalent maximum Loan-to-Value (LTV) ratios offered by British banks to first-time buyers were around 75% during this period, the scheme offers a substantial reduction in the down-payment needed to buy a property. With the government loan covering part of the down-payment, buyers are only required to raise 5% of the property value as a deposit. The explicit goal of the Equity Loan Scheme is that this reduction in the deposit required to the borrower helps households overcome credit constraints.

The Equity Loan Scheme can also help reduce household borrowing costs by reducing interest payments on the combined loan. This occurs via two channels. The first is that no interest or loan fees on the equity loan have to be paid by the borrower for the first five years after the purchase of the house. Subsequently, there is a charge, which depends on the rate of inflation. We calculate the implied subsidy provided through this channel in Section 7. Secondly, by raising the combined deposit to 25%, the equity loan keeps borrowers away from high-LTV-high-interest products that are available in the commercial mortgage market.[5]

The government equity loan can be repaid at any time without penalty. The equity loan is interest-only so, unless borrowers want to sell the property, they do not need to repay the loan at all. When the property is sold, the government will reclaim its 20% equity stake of the sale price. The government thus participates in capital gains and losses.

In our analysis we exploit differences between the English, Welsh and London versions of the Equity Loan Scheme. Regional differences in the scheme are summarized in Table 1. The Welsh scheme was introduced in January 2014 and provided support for the purchase of properties with prices under £300,000.[6] The

[5] This enables households to gain access to more attractive mortgage rates from lenders who participate in the scheme. Eligibility conditions require borrowers to have a suitable credit score and to be able to cover the monthly repayments. Benetton *et al.* (2018) report that lenders adjust mortgage interest rates of HtB borrowers in response to additional default risk associated with lower down payments, but this adjustment is small: it only accounts for 10% of the difference in market interest rates between 75%- and 95%-LTV mortgages.

[6] Scotland also introduced a HtB Equity Loan Scheme in 2013; however, we are not able to exploit the discontinuities at the English/Scottish border. This is because the Scottish Land Registry did not identify new build units until 2018.

London scheme was introduced in February 2016 and offered an equity loan of up to 40% of the unit's price for properties under £600,000 located within the GLA. As we will show in Section 6, the price constraints in both countries, England and Wales, were binding, with substantial bunching of new property sales at these price points emerging after the introduction of the policy.

2.2. Theoretical Framework

In this sub-section we develop a theoretical framework to guide our empirical analysis.[7] Specifically, we develop a simple model of the housing market with heterogeneous households, featuring credit constraints and endogenous housing supply. It is a partial equilibrium model in that we abstract from potential effects of changing credit conditions for new builds on the price of the existing stock.

The framework illustrates how a relaxation of credit conditions affects housing quantities and prices, and how these effects depend on the costs of developing new stock. A relaxation of credit constraints leads to both an increase in prices and an expansion in quantities. Under suitable assumptions – made explicit below – the relative magnitude of the two effects depends on the responsiveness of supply to prices. For low (high) supply responsiveness, the price effect is stronger (weaker) and the quantity effect weaker (stronger). The theoretical insights from this framework can be summarized by the cross-elasticities of quantity and prices taken over the credit conditions parameter and a building cost shifter. We also show that a relaxation of credit conditions, conditional on developers having some market power, can be expected to increase developer profits.

Suppose a two-period economy with a unit mass of households with preferences over a numeraire consumption good c and housing h, as given by a period utility $u(c, h)$ which is continuous, strictly increasing and differentiable in both arguments. Assume in addition that $\lim_{h\to\infty} u(c, h) = \infty$ if $c > 0$ and $u(c, h) > 0\ \forall c, h > 0$. Households enjoy utility at the end of periods 1 and 2, and the discount factor is $\beta>0$.

Households can only obtain a fixed amount of housing consumption $h > 0$ if they buy a new unit and obtain housing consumption normalized to 0 otherwise. We can think of these alternatives as a choice between renting and buying. In this interpretation, this

[7] The model builds on Hilber and Vermeulen (2016) who consider a similar setting but abstract from the role of credit conditions.

formulation is similar to those used in models featuring warm-glow from ownership (Iacoviello and Pavan 2013, Kiyotaki *et al.* 2011, Carozzi 2020).

Households receive an endowment e in period 1 and a location specific income w in period 2, which can be used for consumption or to buy property. Households are heterogeneous in the initial endowment e, which is continuously distributed over the unit interval $[0,1]$ with cumulative density function F_e.

New build units are homogeneous and can be bought in period 1 for an endogenous price P. Credit is available for the purchase of property, yet a minimum down-payment is required corresponding to a fraction $(1-\gamma)$ of the property value. Credit and savings pay interest r. We assume that $w > \frac{\gamma}{1-\gamma}(1+r)$ which ensures that, for sufficiently large h, demand for new build units is determined solely by the credit constraint.[8] Hence, demand is given by the mass of agents that can afford a down-payment $Q_D = 1 - F_e((1-\gamma)P)$. Note that demand is downward sloping as the function F_e is strictly increasing.

There is a unit mass of developable land which can be used to build – at most – a unit mass of housing units. Development costs for new build units depend on local supply conditions and are heterogeneous across land plots. We assume that the development costs are uniformly distributed in the $[0, \nu]$ interval, with $\nu(1-\gamma) > 1$. We assume land is owned by competitive firms which will develop their plot if the price is smaller than or equal to development costs. As a result, the new build inverse supply curve for competitive developers is given by $P = \nu Q$. High values of ν correspond to higher average development costs and, therefore, to a weaker response of quantities to a change in prices. Conversely, low values of ν are associated with a more responsive supply schedule (i.e. a flatter supply curve). We can substitute this expression in demand to obtain an implicit definition for new build equilibrium quantities:

$$Q^* = \left(1 - F_e\big((1-\gamma)\nu Q^*\big)\right) \quad (1)$$

[8] Note that $P \leq \frac{1}{1-\gamma}$. Assumption $w > \frac{\gamma}{1-\gamma}(1+r)$ will therefore ensure that in period 2 all agents are able to pay the remaining part of any loans taken for the purchase of a property, including interest. Large enough h ensures buying property in period 1 is incentive compatible for all households. See the theoretical Appendix.

By differentiating this expression, we can obtain the following four statements regarding the responses of equilibrium prices and quantities to changes in credit conditions (γ), and development costs (ν):

$$\frac{dQ^*}{d\nu} < 0 \qquad \frac{dP^*}{d\nu} > 0 \qquad \frac{dQ^*}{d\gamma} > 0 \qquad \frac{dP^*}{d\gamma} > 0 \tag{2}$$

The first two inequalities indicate that an increase in development costs results in a reduction in equilibrium quantities and an increase in equilibrium prices.[9] The latter two inequalities imply that both quantities and prices respond positively to an expansion of credit. This follows from the increase in demand associated with a credit expansion. The extent to which a change in credit conditions will translate into a change in quantities or prices depends on both the distribution of the initial endowment F_e and the price responsiveness of supply (through ν).

Proposition 1 – The effect of a credit expansion on prices and quantities depends on the distribution of development costs, as measured by ν. Specifically, if e is uniformly distributed and $\nu(1-\gamma) > 1$, then $\frac{dQ}{d\gamma d\nu} < 0 < \frac{\partial P}{d\gamma d\nu}$.

Proof: See theoretical Appendix

Proposition 1 states that, under the specified parameter conditions, the effect of credit expansion on quantities will be smaller, and the effect on prices larger, in high ν markets (i.e., in markets with tighter supply constraints and thus more inelastic long-run supply of housing). This intuition will help us account for regional differences in our estimates of the impact of HtB reported in the next sections.

The assumption of uniform endowments is a sufficient condition, but it is not necessary. Intuitively, this assumption results in linear demand curves. Without linear demand curves we can have that either the first or the second inequality is not satisfied. Hence, the conclusions derived from the uniform case may or may not follow with more general assumptions on the distribution of endowments. This *ex-ante* ambiguity partly motivates the empirical analysis below.

The statements in the derivatives in (2), as well as Proposition 1, are derived for the case of competitive land and housing markets.

[9] See proofs in theoretical Appendix.

Proposition 2 – A credit expansion will result in an increase in total developer profits. That is, the sum of equilibrium profits across developers $\Pi(\mathrm{P})$ is increasing in γ.

Proof: See theoretical Appendix

This result hinges on the assumption that developers own all land, preventing entry from other firms from eroding profits. The notion that developers have some market power is reasonable in our case, as the residential construction market is characterized by substantial concentration and high returns. We test empirically whether Proposition 2 is satisfied in Section 4.4.

3. Data and Descriptive Statistics

3.1. Main Data Sources

Our empirical analysis employs geo-located data on housing sales in England and Wales, including information on unit characteristics and transaction prices. Our main data source is the Land Registry Price Paid Data (or short 'Land Registry'), which covers most residential and all new build residential transactions in England and Wales. The dataset includes property sales from 2010 to 2019, recording the transaction price, postcode, address, the date the sale was registered (which proxies for the transaction date), and categorical data on dwelling type (detached, semi-detached, flat or terrace), tenure (freehold or leasehold) and whether the home is a new build property. We use the National Statistics Postcode Lookup Directory to match properties in the dataset to coordinates and wards.

Between 2010 and 2019, the Land Registry recorded 948,553 sales of new housing units. We use these as a proxy for construction activity. All sales are geo-coded using address postcodes. In our spatial discontinuity analysis, we use new build transactions taking place near the GLA boundary and near the English/Welsh border. Specifically, we select all new build transactions within 5km from the GLA boundary and within 10km from the English/Welsh border after removing a small set of observations that we identify as being sold in bulk between developers.[10] We use a 10km bandwidth for

[10] The number of transactions for the resulting samples are reported in Appendix Table A1. We exclude a total of 1041 sales. These are transfers to non-private individuals, in bulk of over 20 sales within the same building and month. We exclude these from our sample as they likely correspond to transfers between companies within a conglomerate and thus do not represent genuine market transactions. This exclusion does not affect any of the main results in the paper.

the latter exercise because transactions near the English/Welsh border are sparser. We also use areas near the Greater Manchester boundary in a placebo test.[11]

In addition, we use Energy Performance Certificate (EPC) data that contains information on the floor area and other physical characteristics of newly built units. We match this data to the Land Registry to augment the latter dataset with additional information on the transacted newly built units.[12] Demographic neighbourhood characteristics at ward level are collected from the 2011 Census. We include the fraction of married residents, and the fraction of residents with level-4 and above educational qualifications as controls in some specifications.

3.2. Descriptive Statistics

In Panel A of Table 2, we present summary statistics for the sample of new build sales located within 5 kilometres of the GLA boundary taking place between January 2010 to December 2019. There are 41,357 newly built property transactions in this sample. The average house price is £389,440, and the average size of these properties is 86.8 square meters. Panel B of Table 2 shows the descriptive statistics for the baseline sample of new build transactions within 10 kilometres of the English/Welsh border taking place between 2010 to 2019. The average house price in this region is £232,536, and the average size of these properties is 101.4 square meters.

When estimating the effect of the policy on housing construction, we assemble a ward by month panel using data from January 2010 to December 2019. We obtain ward-level observations by aggregating from individual new build sales. Panels C and D of Table 2 document the descriptive statistics of our estimation sample for the construction effect. The datasets for the GLA boundary-area and the English/Welsh

[11] Greater Manchester is the second largest travel to work area in the United Kingdom and arguably the one most comparable to London.

[12] EPCs provide information on buildings that consumers plan to purchase or rent. Since 2007 an EPC has been required whenever a home is constructed or marketed for social rent, private rent or sale. We use a dataset that contains all EPCs issued between 2008 and 2019. The dataset includes the type of transaction that triggered the EPC, the energy performance of properties and their physical characteristics. Following Koster and Pinchbeck (2017), we merge the EPC data into the Land Registry using a sequential match strategy. First, we match a Land Registry sale to certificates using the primary address object name (PAON; typically, the house number or name), secondary address object name (SAON; typically, the identification of separate unit/flat), street name, and full postcode. We then retain the certificate that is closest in days to the sale or take the median value of characteristics where there is more than one EPC in the same year as the sale. We repeat this exercise for unmatched properties but allow one of the PAONs or SAONs to be different. Our final round of matching is on the full postcode. The matched dataset provides us total floor area; whether the property has a fireplace or not; total energy consumption and total CO_2 emission of the property.

border-area consist of 411 wards and 204 wards, respectively. The propensity to have at least one new build transaction in any month and ward is 0.2 for both the GLA sample and the English/Welsh sample. On average, 0.8 new units are built each month in a ward near the GLA boundary and 0.5 near the English/Welsh border.

To conduct our analysis of developer performance, we construct a developer/construction company panel that covers 78 companies during the period 2010 to 2019. The panel includes financial information of these companies from Orbis. It also includes information on whether the companies are registered with a HtB agency. A builder must be registered with one of the regional government offices managing the scheme for its properties to be eligible for a HtB equity loan. The full sample of developers is our *Difference-in-Differences* sample. It is obtained after combining a list of the main builders in the United Kingdom from Zoopla – one of the main property websites in the country – and financial data from Orbis. This list includes residential developers, commercial developers and construction companies. We present the descriptive statistics for this sample in Panel A of Table 3. The average turnover of these companies is £ 540 million, and the average net profit before tax is £ 64 million. In addition, we include hand-coded data on the fraction of properties sold through the scheme from annual reports in a selected sample of 30 residential developers. This is our *intensity* sample, and Panel B of Table 3 documents its descriptive statistics. Finally, to mitigate the concern of more profit-driven developers (with different characteristics) self-selecting into the HtB-scheme, we compute the ratio of HtB completions relative to the number of total new build transactions at the NUTS-1 level as an instrument for the observed HtB-status of developers and link this local HtB-intensity to developers using their headquarters' address information. This is our *HtB completion* sample and covers 69 companies. We report the summary statistics for this sample in Panel C of Table 3.

3.3. Local Supply Conditions

Below, we report separate estimates of the impact of the generosity of HtB schemes obtained from a sample of properties near the GLA boundary, and a sample of properties near the English/Welsh border. We choose these two areas because they both provide an ideal quasi-experimental setting to identify the economic effects of HtB.

One crucial difference between our two focal areas – the area near the GLA boundary and the area near the English/Welsh border – is that the former has overall vastly less responsive supply, driven by both, tighter local planning regulations and a relative scarcity of undeveloped developable land (Hilber and Vermeulen 2016). As shown above, theory suggests that the positive impact of HtB on house prices should be much larger – and the positive impact on new construction much smaller – in the area near the GLA boundary.

To illustrate the differences in supply conditions between the areas, we employ a number of measures that capture long-term housing supply constraints. These measures are the share of land designated as green belt (provided by the Ministry of Housing, Communities and Local Government), the average planning application refusal rate taken over the period from 1979 to 2008, the average share of developed developable land, and the average elevation range (all derived from Hilber and Vermeulen, 2016). We calculate these measures for the areas employed in our analysis using Local Planning Authority (LPA)-level data and LPA surface areas as weights.[13] We also report similar descriptive statistics for LPAs around the Manchester boundary, as we use this region in a placebo test (reported in Section 5).

Table 4 (rows 1 to 4) illustrates the differences in supply conditions between the three areas. The most striking difference between the two focal areas lies in the share of 'green belt' land. Land in green belts is typically off limits for any development (residential or commercial) and thus represents a 'horizontal' supply constraint. This share is 66.5% for boroughs along the boundary of the GLA but only 3.8% for English boroughs along the English/Welsh border. Another measure to capture physical supply constraints is the share of developable land already developed. This share is 27.6% for boroughs along the GLA boundary (with developable land mostly being green belt) but only 6.3% for English boroughs along the English/Welsh border.

The arguably quantitatively most important long-term supply constraint are restrictions imposed by the British planning system (Hilber and Vermeulen 2016). The weighted average of the planning application refusal rate is 35.6% for boroughs along the GLA boundary and 27.2% for English boroughs along the English/Welsh border.

[13] We do not currently have data for LPAs on the Welsh side of the English/Welsh border. We expect that the differences between the GLA and the English/Welsh border area would be even more striking when taking account of the data from the Welsh LPAs.

While the area near the English/Welsh border is subject to greater topographical (slope related) supply constraints, Hilber and Vermeulen (2016) demonstrate that these constraints, while statistically significant, are quantitatively unimportant in explaining local price-earnings elasticities.

Lastly, it is important to point out that the area near the GLA boundary is not only characterized by vastly more restrictive supply conditions, but these constraints are also significantly more binding in practice, simply because aggregate housing demand there is much stronger. To illustrate this point, consider a ten-story height restriction in the heart of a superstar city such as London and compare it to the same constraint in the desert. The restriction is extremely binding in the former location, while completely irrelevant in the latter.

To explore the differences in supply responsiveness across the three areas further, we employ the estimated coefficients from Hilber and Vermeulen (2016) to compute an implied house price-earnings elasticity. Table 4 (rows 5 and 6) reports our estimated elasticities based on these coefficients. Using the OLS estimates, we find that the price-earnings elasticity along the GLA boundary (0.40) is higher than that of the area along the Greater Manchester boundary (0.28), which in turn is higher than the elasticity near the English/Welsh border (0.25). As two of the three supply constraints measures employed in their estimation, refusal rate and share developed land, are likely endogenous, we employ the instrumental variable strategy proposed in Hilber and Vermeulen (2016). This provides exogenous variation in our supply constraint measures, which we use to re-compute the unbiased price-earnings elasticities. The rank order remains unchanged. The GLA has again the highest elasticity (0.21) followed by Greater Manchester (0.16) and the English/Welsh border area (0.13).

The higher price-earnings elasticity along the GLA boundary suggests that housing prices respond more strongly to a given change in local housing demand. This also implies a lower supply price elasticity in the GLA boundary area. In the next section, we outline our identification strategy and discuss how we measure the impact of HtB on house prices, construction activity and the financial performance of developers.

4. Main Empirical Analysis

Our empirical strategy is designed to test the impact of HtB on house prices and housing construction by exploiting spatial differences in the intensity of the HtB policy.

HtB Wales – rolled out nine months later than in England – offered a government-backed loan for the purchase of new build properties under £300,000, compared to £600,000 in England. Thus, the policy was more generous on the English side of the boundary. There were also differences in the intensity of the HtB policy between the GLA and its surroundings, starting in 2016. In this case, the difference lies in the size of the government-backed loan available to households. London-HtB offers loans of up to 40% of a new build's value, while this figure is 20% outside the GLA boundary. We exploit these regional policy differences in a Difference-in-Discontinuities design combining time variation in prices and new build construction with local variation in policy intensity around the regional boundaries.

The samples of new build properties used in the analyses of prices and construction effects near the GLA boundary and the English/Welsh border are illustrated in Panels A and B of Figure 1, respectively.[14] Our boundary approach is meant to ensure that we are comparing properties affected by similar economic and amenity shocks, as compared to a standard Difference-in-Differences strategy that simply takes whole regions as treatment and control groups. The identifying assumption in both cases can be likened to the typical assumption of parallel trends: in the absence of the policy, prices and construction on either side of the boundary would have followed a parallel evolution over time.

4.1. Graphical Illustration

Figure 2 depicts the evolution of house prices on both sides of the GLA boundary and English/Welsh border, respectively, and indicates that prices moved in parallel prior to the implementation of the policy.[15] Panel A shows that the gap between house prices inside and outside of the GLA starts to widen only after the introduction of London's Help to Buy scheme. The gap appears in 2017, in line with the fact that the proportion of HtB sponsored purchases in London grows substantially during that year.[16] Panel

[14] Appendix Figure B1 depicts the corresponding map for our placebo sample of new build sales near the Greater Manchester boundary.
[15] The price index is constructed by estimating a linear regression of log prices on property characteristics (property type dummies for detached, semi-detached and terraced properties, log property size, a leasehold dummy, measures of energy efficiency) and postcode fixed effects. The lines in Panels A and B of Figure 2 correspond to time dummies included in that specification.
[16] The proportion of HtB completions relative to the number of new build transactions in London increases from 10.7% in 2015 and 12.7% in 2016, to 24.5 % and 31.9% in 2017 and 2018, respectively. This is likely because it takes time for developers and home buyers to learn about and adjust to the new HtB scheme in London.

B, in contrast, shows no substantial divergence between price indices on both sides of the English/Welsh border.

Figure 3 depicts the average number of units built by ward at the GLA boundary and the English/Welsh border, respectively. Again, we see that the evolution of building activity followed reasonably parallel trends prior to the implementation of the policy. Panel A shows no substantial divergence between housing construction inside and outside of the GLA after the implementation of HtB, while Panel B reveals that a gap emerged in the building activity on both sides of the English/Welsh border after the policy was introduced, indicating that the more generous English scheme stimulates construction at the English/Welsh border where supply is less constrained.

We then provide a series of graphs in Figures 4 to 7 that illustrate our main results. Figure 4 depicts the prices for newly built units at different distances from the GLA boundary. Positive distances correspond to locations inside the GLA, while negative distances refer to locations outside of this area. Circles depict the mean value of new build house prices for 500-meter-wide distance bins with the size of each circle being proportional to the number of observations in that bin. Lines in both panels represent fitted values estimated separately on each side of the boundary. Gray bands around them represent 95% confidence intervals.[17] Panels A and B illustrate results before and after the introduction of London's HtB, respectively. Comparing both panels, we find that a discontinuity in prices at the boundary emerges after the implementation of London's HtB. We interpret this as evidence that differences in the size of available equity loans at the boundary led to a significant and positive effect on the price of newly built properties within London. We test this formally in Section 4.2.

Figure 5 illustrates our results for the new build price effect at the English/Welsh border. Circles depict the mean value of house prices for 1000-meter-wide distance bins. As above, solid lines represent fitted values estimated on both sides of the boundary. In this case, however, we do not observe a spatial discontinuity of house prices in either Panel A or B. Hence, the difference in the scheme at the border did not generate an appreciable difference in new build prices.

We conduct a similar exercise looking at changes in construction volumes at these boundaries before and after the corresponding changes in HtB. Results are illustrated

[17] Appendix Figures B2 and B3 report results when using second order polynomials on either side of the boundary.

in Figures 6 and 7. The former shows construction as measured by new build sales near the GLA boundary with Panels A and B, respectively, corresponding to the periods prior and post implementation of London's HtB. We do not find a spatial discontinuity in homebuilding at the London boundary in either period. Figure 7 depicts results for the English/Welsh border before and after the English HtB policy was implemented. In this case, we find a wider discontinuity in Panel B, indicating more building took place on the English side of the boundary after the policy was introduced.

Overall, these graphs indicate that more generous versions of the HtB policy triggered a price but no quantity response in the supply inelastic areas around London. Conversely, the policy generated a quantity but no price response in the relatively supply elastic areas around the English/Welsh border. This is in line with intuition and with our theoretical framework, which suggests that price and quantity responses to shifts in demand depend on the shape of the supply curve. Below, we document reduced-form estimates for the magnitudes of these effects.

4.2. The Impact of Help to Buy on House Prices

We explore the impact of differences in the generosity of the HtB-schemes at the London boundary and at the English/Welsh border on prices P by employing a Difference-in-Discontinuities framework. We estimate the following equation:

$$\ln\left(P_{ipjt}\right) = \phi_p + \beta HtB_i \times Post_t + \delta_t + \gamma' X_i + \tau' Z_j \times d_y + \gamma_y Distance_i \times d_y + \varepsilon_{ipjt} \qquad (3)$$

where i indexes individual properties, p indexes the postcode, j indexes the (ward-level) neighbourhood, t indexes the month, and y indexes the year. The variable HtB_i is a dummy that takes value 1 in the region with a more generous HtB policy (i.e. inside the GLA or on the English side of the English/Welsh border). The variable $Post_t$ represents a dummy taking value 1 if individual transaction i occurs after the change in policy (e.g. London's HtB was introduced in February 2016, so $Post_t$ takes value 1 from March of that year). A vector of postcode fixed effects, represented by ϕ_p , is included to account for fixed differences in amenities and other local characteristics across locations. Likewise, we include a set of individual housing characteristics X_i to account for differences in the attributes of sold units.[18] We include a set of (year-month) time dummies δ_t to account for aggregate changes in prices in each sample. A vector of neighbourhood characteristics Z_j interacted with year dummies d_y are included to account for time-varying changes in neighbourhood characteristics unrelated to HtB. Finally, we include the distance to the boundary interacted with year dummies d_y to account for potential time varying shocks that differ spatially.[19]

Our parameter of interest is β. It measures the effect of differences in the intensity of the HtB policy on the price of new build properties.

We estimate this equation by OLS on new build properties, clustering standard errors at the postcode-level to account for potential spatial correlation in local price shocks. In the case of the London HtB scheme, we use a 5km bandwidth around the GLA boundary. We use a 10km bandwidth around the English/Welsh border. In the robustness checks section, we show that our results are robust to alternative bandwidth choices.

Table 5 presents results obtained from estimating equation (3) using the sample of transactions around the GLA boundary. We include different sets of covariates sequentially from columns 1 to 5. Column 1 controls for time effects and independent linear terms in distance of each property to the GLA boundary. Column 2 adds a vector of housing characteristics such as type of the property, energy consumption, and tenure (freehold vs. leasehold). Column 3 adds postcode fixed effects. In column 4 we include neighbourhood characteristics from the census interacted with year effects. Finally, in column 5, we allow for heterogeneous spatial price trends by controlling for interactions between distance from the GLA boundary and year dummies. Our preferred specification is column 5.

The resulting estimates show that London's HtB policy increased the price of newly built houses inside the GLA by between 5% and 8% depending on the specification, with 4 out of 5 estimates being significant at the 5% level. The average property price in this sample is £389,440, suggesting that homebuyers are paying roughly £30k more

[18] These controls are included to account for the fact that the policy may induce a change in the characteristics or the location of the units built by the developers. We return to this in Section 6.2.
[19] In an alternative specification, we omit the postcode fixed effects and control flexibly for distance to the boundary by estimating different linear terms in the distance, specified separately on either side.

to buy newly built properties inside the GLA because of London's HtB (compared to the less generous English-version of the scheme). In Section 7, we compare this amount to the implicit interest subsidy provided by the equity loan granted by the scheme.

Table 6 summarizes the results from estimating equation (3) for the sample of properties around the English/Welsh border. Again, we successively include additional controls from columns 1 to 5. Once we control for postcode fixed effects, we observe no significant effect of the policy on the price of newly built properties.[20] The point estimates in columns 3 to 5 are positive but small, ranging between 0.1 to 0.9%, and not statistically significant.

4.3. The Impact of Help to Buy on Housing Construction

The government's equity loan is available only for the purchase of newly built units. In this way, the government attempts to ensure the policy triggers additional residential construction. To formally test whether this is the case, we estimate the effect of differences in the intensity of the policy on construction activity. We again use a Difference-in-Discontinuities specification, diverging from the one employed to study price effects in that we obtain our estimates using a ward level panel built by aggregating new build counts at the ward level for every month. We estimate:

$$New\ builds_{jt} = \omega_j + \beta HtB_j \times Post_{t-12} + \delta_t + \tau' Z_j \times d_y + \gamma_y Distance_j \times d_y + \varepsilon_{jt} \quad (4)$$

where j indexes wards, t indexes months, and y indexes years. The dependent variable is now $New\ builds_{jt}$, which can represent either the number of new build transactions in ward j and period t, or a dummy taking value 1 if there are any new build sales in ward j and period t. As above, the variable HtB_j is a dummy taking value 1 in the area with a more generous HtB policy. The variable $Post_{t-12}$ represents a dummy that takes value 1 if transactions in ward j occur after the difference in policy arises. The variable is lagged by twelve months to account for the likely delayed response of construction to the policy shock.[21] We include a set of ward fixed effects, represented

[20] This finding does not depend on the ordering of introducing controls. If we introduce census variables-by-year controls in our column 2 specification instead of postcode fixed effects, we also obtain a small and insignificant coefficient.
[21] Construction lags in the UK tend to be long by international standards, often in excess of 12 months. As a robustness check, we estimate a contemporaneous specification.

by ω_j and time fixed effects δ_t.[22] Z_j are neighbourhood characteristics from the 2011 Census interacted with year dummies d_y. In addition to controlling for ward fixed effects, we include the distance to the boundary interacted with year dummies to account for potential time varying shocks that differ spatially. In all specifications we cluster standard errors at the ward level.

Our parameter of interest is β, which measures the effect of differences in the intensity of HtB on new construction. The differences in intensity are not the same across the English/Welsh border and across the GLA boundary. Therefore, we obtain separate estimates for these two exercises.

Table 7 summarizes the results from estimating equation (4) for the sample including wards around the GLA boundary. We define the post-HtB period as extending from February 2017 to December 2019 – starting one year after the implementation of London's HtB – to allow for a one-year construction lag. Table 7 reveals that London's HtB had neither a significant effect on construction volumes nor on the probability that any newly built property was sold in a ward. Coefficients are insignificant and small in all specifications, indicating that the increase in the size of available equity loans inside the GLA-boundary did not lead to an increase in housing construction.

In Table 8, we provide estimates of equation (4) for wards around the English/Welsh border. As above, the post-treatment period is defined as starting one year after the introduction of the English HtB-scheme. We find a significant and positive effect of HtB on housing construction in all specifications. Our estimates suggest that the higher eligibility threshold on the English side of the border increased the number of new build transactions at each ward by 0.4 on average, and the propensity for any new build construction at each ward by about 8%. These results are consistent with the predictions from our theoretical framework that indicate that HtB has differential effects in London and the areas around Wales as a consequence of differences in supply conditions in the two areas.

4.4. The Impact of Help to Buy on the Financial Performance of Developers

Uncovering how HtB affected the financial performance of developers can help us identify some of the beneficiaries of the policy. Theoretically, the HtB policy can be

[22] We also provide estimates that are obtained by controlling flexibly for the distance to the boundary, omitting ward fixed effects.

expected to induce an increase in revenue of existing developers participating in the scheme.[23] Moreover, as stipulated in Proposition 2, barriers to entry and imperfect competition in housing production and land markets imply that the policy should translate into increased profits. This, however, hinges on the increase in revenue not being fully offset by an increase in the cost of land after the implementation of the policy.[24]

Lack of detailed information on the location of developers' assets prevents us from deploying the spatial techniques used in our analysis of price and construction effects. To nevertheless study the effects of HtB on the financial performance of developers empirically, we instead employ a Difference-in-Differences strategy and use our developer dataset, covering 78 large British developers and construction companies.[25] The dataset includes information on the developers' financial performance and, crucially, on the participation of these firms in HtB. We use our dataset to compare how the change in the performance of firms before and after 2013 varied with their participation in the scheme. For this purpose, we estimate a fixed effect model specified as:

$$Fin\,performance_{kt} = \beta HtB_k \times Post_t + \alpha_k + \delta_t + \varepsilon_{kt} \quad (5)$$

$Fin\,performance_{kt}$ is an indicator of various measures of financial performance for developer k in year t. We look at turnover (i.e. total revenues), gross profits, and net profits before taxes. The measure HtB_k captures a developer's engagement with the policy. The variable $Post_t$ takes value 1 in 2013 and in subsequent years. Finally, α_k is a developer fixed effect and δ_t represents a set of year dummies.

Estimates of β will capture the impact of the policy on developers' financial performance measures under the assumption that unobservables ε_{kt} are uncorrelated with $HtB_k \times Post_t$ conditional on individual and year fixed effects.

To ensure the internal validity of our Difference-in-Differences model we first visually inspect the crucial parallel trend assumption. In Figure 8, we plot our three yearly average financial performance indicators – adjusted for individual company fixed-

[23] The increased supply could in principle be taken up exclusively by new entrants. Yet the presence of economies of scale in housing production and the learning curve required to navigate the British planning system mean that the volume of new entrants will probably be very small.

[24] In our model, this is ruled out because land is owned by developers, so land rents are included in profits.

[25] Our regression samples only cover a small number of relatively large developers and are thus only partially representative of the entire industry.

effects – for the HtB and non-HtB groups of developers before and after the policy. The pre-trends are reasonably parallel, and we observe a clear divergence after 2013, with substantial growth for developers registered for HtB. The plots in Figure 8 thus do not only provide support for the parallel trend assumption, they are also consistent with the notion that developers improved their financial performance because of HtB. An additional implication is that, on the supply side of the residential market, the benefits of the scheme may not have gone exclusively to landowners.

Now turning to our estimates of equation (5), we use two alternative measures to capture a developer's engagement with HtB. The first is based on the registry of developers in regional HtB offices across the country. In this case, the variable HtB_k is a dummy taking value 1 if the developer is included in the HtB-registry. We can estimate this specification for our full Difference-in-Differences sample of 78 developers. Our second measure of HtB_k is based on detailed information on the fraction of the units produced by developer k that were sold under the HtB scheme, averaged over our sample period. Because this information is only available for a subset of companies, we can only estimate this specification with our reduced 'intensity sample' covering 30 developers.

Table 9 presents our estimates for the effect of the scheme on revenues, gross profits and net profits before taxes. Estimates in Panel A indicate that participation in HtB – as measured with our registration dummy – increases revenues substantially, with HtB-participants obtaining over 57% higher revenues compared to non-participants.[26] Coefficients for gross and net profits are even larger, indicating a large effect on developer performance.

Panel B presents estimates of the effects for our continuous measure of HtB participation using our intensity sample. The first column shows that a 1 percentage-point increase in the fraction of HtB-properties supplied by a developer, leads to a roughly 1% increase in revenue. The effect is large and significant. The estimates for gross profits and net profits, displayed in columns 2 and 3 are even larger, suggesting that changes in costs – e.g. costs of acquiring land – did not offset the changes in revenue. Again, these estimates suggest that the policy improved the financial performance of residential developers.

[26] The coefficient β is 0.45, so we can write the proportional difference in revenues as $\Delta r = e^{0.45} - 1$.

Some caution is warranted when interpreting the findings in Panels A and B of Table 9. There are substantial observable differences in characteristics between the developers self-selecting into the HtB-scheme and other developers in the sample. For example, luxury developers typically are in the control group, as they will not normally be registered with HtB. We can only give our estimates a causal interpretation if these differences have a time-invariant influence on the financial performance of developers. While we would argue that this assumption is plausible, we address the concern of possible self-selection by employing an Instrumental Variable approach.

Our instrument for a developer *k*'s engagement in HtB is the ratio of HtB completions in the NUTS-1 region, in which developer *k* has its headquarters, relative to all completions in that region. This measure can be expected to be strongly correlated with developer *k*'s engagement in HtB. However, it is not directly affected by developer *k*'s characteristics. As such the instrument helps us address the concern that inherently more profit-driven developers may self-select into the HtB-scheme and that this may explain the better financial performance of these developers, rather than HtB itself increasing developer returns and profits.[27]

Panel C of Table 9 reports second stage estimates of the effect of HtB-registration on developer performance, using our local HtB exposure-instrument in the first stage. The resulting second-stage estimates are consistent with the Difference-in-Differences estimates reported in Panel A and are statistically significant at least at the 5% level. Panel D of Table 9 reports the corresponding first-stage result, indicating that developers headquartered in regions that have become more HtB-intensive over time have become more likely to participate in the scheme. This instrument is however rather weak – with an F-statistic of 4.8 – so we must interpret our IV results with caution.[28] It is worth noting in this context, however, that results are qualitatively similar when we use the local HtB exposure measure directly rather than as an instrument – see Panel E of Table 9.

[27] The identifying assumption is that developers are more likely to become engaged in HtB in local markets in which demand for HtB-equity loans has become stronger over time and that, conditional on developer and year fixed effects, the uptake in HtB loans in a local area only affected developer *k*'s financial performance via affecting developer *k*'s propensity to become engaged in HtB.

[28] Following Andrews et al. (2018) and Sun (2019), we also report the Anderson-Rubin 90% confidence interval for our estimates in Panel C of Table 9.

While the various estimates reported in Table 9, individually, must be interpreted with caution, taken together, these results clearly point to a causal positive effect of HtB on the financial performance of participating developers.

5. Robustness Checks

We now turn to a series of robustness checks and placebo tests to confirm our main findings and provide additional validation to our research strategy. First, we look at whether our findings can be explained by displacement of demand across the English/Welsh and London boundaries. Second, we provide estimates employing displaced versions of the boundaries used in the main analysis and the area around the Greater Manchester boundary to construct suitable placebos. A final set of tests considers robustness to varying chosen bandwidths or the period used in the analysis.

5.1. Displacement of Homebuyer Demand Across Boundaries

The potential displacement of demand across either the GLA boundary or the English/Welsh border is an important threat to our identification strategy. Displacement could occur if the policy induces short-distance sorting of prospective buyers so that, for example, demand for housing falls outside the GLA boundary as a result of the policy. This would violate the stable unit treatment value assumption required to interpret our quantitative estimates of the price or construction effects as the outcome of the policy. Fortunately, if we assume demand displacement is relatively local – i.e. occurring mostly over short distances – we can use our samples to evaluate whether this is indeed a problem and what is its impact on each set of estimates. We do so with two different strategies.

The first is to reproduce our Difference-in-Discontinuities baseline estimates, sequentially dropping the transactions closest to the boundary. If the displacement of demand across the boundary of interest is important and happens over relatively short distances, then excluding observations next to the boundary should partially correct our estimates for demand sorting. Estimates for the price effect of London's HtB obtained after excluding different bands around the GLA boundary are reported in Panel A of Appendix Figure B4. We observe that we can exclude transactions taking place within up to at least 2km of the GLA boundary – 40% of the bandwidth – without a significant effect on our estimates. In Panel B, we report an analogous figure for the area around the English/Welsh border. Again, excluding observations within 40% of the bandwidth around the border does not affect the conclusion that differences in the HtB scheme's generosity did not lead to a significant price effect.

It is also possible that our housing construction estimates are biased by demand sorting. To explore this, we obtain estimates after iteratively excluding newly built properties around the boundary. In Panel A of Appendix Figure B5, we report the coefficients that capture the impact of HtB on construction near the GLA boundary for the different sample restrictions. All coefficients are statistically indistinguishable from zero. We provide results for the construction effect of HtB near the English/Welsh border, again for different sample restrictions, in Panel B. The estimated coefficients are fairly stable between 0.3 and 0.4 and the point estimates are substantially larger in absolute value compared to the coefficients estimated for the area near the GLA boundary.

The second strategy seeks to directly test whether there is any evidence of displacement across our boundaries of interest. To do so, we follow the intuition in Turner *et al.* (2014) and compare transaction prices close to and far away from the boundary *within* each side.[29]

We focus on the statistically significant effects of HtB, that is, the price effect at the GLA boundary and the construction effect at the English/Welsh border.

When looking at the price effect in London and the role of demand displacement around the GLA boundary, we split the sample into two sub-samples corresponding to property sales on each side of the boundary. The displacement hypothesis has specific predictions regarding how demand changes *within* each spatial band around the border. In the case of the sub-sample of properties *inside* of the GLA, a local displacement of demand would result in an increase in new build prices close to the boundary relative to prices further inside the region. Conversely, for sales *outside* of the London region, displacement of demand would reduce prices close to the boundary relative to prices further out into the periphery. These predictions are easily testable using a modified version of equation (3) in which we replace $HTB_i \times Post_t$ with $close_i \times Post_t$, where $close_i$ is a dummy taking value 1 for properties within 2.5 of the boundary. Estimates for London, split by sub-sample, are reported in the first and third column of Appendix Table A2. The results are not consistent with evidence of displacement. The estimates for $close_i \times Post_t$ are insignificant and small in both sub-samples.

[29] Specifically, we follow the approach in the external effect regressions in Section 2.5 of Turner *et al.* (2014).

Statistical power is quite low in these sub-samples, partly because we are using a binary variable to capture distance. To avoid this, we can use another modified version of equation (3) but now replacing $HtB_i \times Post_t$ with $dist_i \times Post_t$ where $dist_i$ is a linear term in distance to the GLA boundary. Estimates for the coefficient on $dist_i \times Post_t$ for each sub-sample are provided in columns 2 and 4 of Appendix Table A2. The coefficient is insignificant and small outside of London. Importantly, the estimate is not negative inside of London, as displacement would predict in this case. We conclude from these analyses that local displacement of demand across the London boundary is negligible.

When considering the statistically significant construction effect at the English/Welsh border, we can test directly for evidence of displacement by estimating a version of equation (4) in which $HtB_j \times Post_{t-12}$ is replaced with $close_j \times Post_{t-12}$ where $close_j$ is a dummy taking value 1 for wards with centroids within 5km of the boundary. We estimate this separately for the English and Welsh sub-samples around the border with the dependent variable being a dummy that takes the value 1 if there was any sale of new build properties in that ward-month pair. The results, which we report in columns 1 and 3 of Appendix Table A3, are not consistent with the pattern that would emerge if displacement of demand across the boundary was significant (both coefficients are small and insignificant).[30] We report a similar analysis using a linear term for distance in the interaction for completeness in columns 2 and 4 of Appendix Table A3. These results confirm the earlier finding.

Collectively, these estimates indicate local demand displacement is either not taking place or is negligible. As a result, we believe this is unlikely to induce a substantial bias in our baseline estimates of the effects of HtB.

5.2. Placebo Tests

We consider two types of placebo tests for our analysis of the effect of differences in policy intensity on the price of new build properties and construction. First, we replicate our estimates for both outcomes (price and construction) using sales of units within 5km of the boundary of the *Greater Manchester* area. No specific HtB-scheme was put in place in this area, so the eligibility conditions and the maximum size of the

[30] Displacement would result in a negative coefficient for $close_j \times Post_{t-12}$ for the Welsh subsample and a positive coefficient for the English subsample. Both estimated coefficients are insignificant and positive in Appendix Table A3.

loan do not vary at this boundary. Estimates for the Manchester placebo for prices and construction are reported in Appendix Tables A4 and A5, respectively. Both tables indicate no statistically significant effects at the boundary, as expected.

Second, we displace the boundaries – i.e. the GLA boundary and the English/Welsh border – relative to their initial positions. In the case of the GLA boundary, we report three estimates per outcome: one obtained by keeping the initial boundary in place (this simply reproduces the headline estimates at the GLA boundary), one obtained by displacing the boundary 5km closer to the centre of London, and, finally, one obtained by displacing the boundary 5km further out. The distances of these displacements ensure that we only use observations on one side of the boundary in each of the placebos. We conduct a similar analysis for the English/Welsh border, displacing the boundary by 10km either into Wales or into England. Results for these placebos, alongside our main estimates, for both price and construction effects, are illustrated in Appendix Figure B6 (Panels A to D). All estimates obtained for displaced boundaries are statistically insignificant, as expected.

5.3. Other Robustness Checks

Our main estimates are obtained using observations within 5km of the GLA boundary and 10km around the English/Welsh border. Results with alternative bandwidths for each exercise are reported in columns 1 and 2 of Appendix Tables A6 and A7, respectively, and indicate no substantial difference in the magnitude of the estimated effects of interest, indicating that our results are robust to reasonable bandwidth choices.

Our construction estimates in Tables 7 and 8 allow for a one-year construction lag to incorporate the fact that demand cannot instantaneously translate into more new build sales. Our results are robust to this choice of timing. Columns 3 and 4 of Appendix Table A6 and columns 4 and 6 of Appendix Table A7 report estimates of contemporaneous construction effects (i.e., the post-treatment-period is defined as the implementation date of the policy) for the London boundary and the English/Welsh border, respectively. Again, we find that HtB does not have a significant impact on housing construction at the GLA boundary but increases construction significantly at the English/Welsh border.

Finally, we test whether our findings are robust to varying the period used in the analysis around the English/Welsh border. The English version of HtB was implemented 9 months before the Welsh version was introduced. Thus, the estimated effects obtained in Tables 6 and 8 have to be interpreted as weighted averages of the impact of the different eligibility conditions of HtB at the border (i.e., the fact that the price threshold on the English side of the border is twice that in Wales) and differences arising from the timing of implementation in both locations. To cleanly identify the effect of the different eligibility conditions, we drop observations between April and December 2013 (i.e., the time period with only English HtB) and replicate our main estimates with this subsample. Results are reported in columns 3, 5 and 7 of Appendix Table A7. The estimated price effect at the English/Welsh border continues to be statistically insignificant, while the estimated construction effect continues to be positive and significant.

6. Additional Results: Bunching and Housing Size

6.1. Price Bunching

The English HtB policy is only available for properties purchased under 600,000 GBP. As a result, the English HtB program led to significant bunching of sales right below the price threshold. Appendix Figure B7 shows two histograms of new build frequencies for prices between £550,000 and £650,000. The left Panel A represents properties sold in the period from January 2012 to March 2013, before the implementation of HtB in England. The right Panel B corresponds to a histogram for properties sold between April 2013 and December 2019, after HtB was introduced. We observe a substantial increase in bunching in the price distribution of new builds just below £600,000. Appendix Figure B8 represents similar histograms for Wales before and after the introduction of its own version of HtB for prices between £250,000 and £350,000. We can see that the introduction of HtB also led to bunching, albeit somewhat less pronounced, of new build sales just below the corresponding threshold – in this case corresponding to £300,000. The fact that bunching is also observed in Wales is important because it shows that the £300,000 threshold induces a change in market outcomes, and a local increase in demand. It therefore motivates the strategy used to measure price and quantity effects at the English/Welsh border.

One issue to consider when identifying the degree of price bunching is round-number bunching. As shown in Appendix Figures B7 and B8, there was already some bunching at the £600,000 and £300,000 thresholds before the policy was in place. To deal with this issue, we employ three strategies. We first show the evolution of newbuild sale volumes by price bins over time. In Appendix Figure B9 we group sales for England into £10,000 price bins and then plot the evolution of the fraction of new builds over total sales for each bin from 2010 to 2019. The black line represents the price bin of interest, £590,000 to £600,000. Grey lines correspond to the other bins between £510,000 and £700,000. We can see that a gap between the black and the grey line appears in 2014 and widens substantially from 2015, implying a significant amount of bunching of new builds at £600,000 after this year, conditional on round-number bunching in the price distribution of all sales.

Second, we use the total number of sales – of newbuilds *and* existing units – to normalize for a baseline level of round-number bunching. Appendix Figure B10 shows the fraction of new builds over total sales for England and for £5000 price bins averaged over the period between April 2013 and December 2019. Horizontal dashed lines represent averages above and below the £600,000 threshold. We observe significant bunching at £600,000. Appendix Figure B11 repeats this exercise for Wales. Horizontal dashed lines represent averages above and below the £300,000 threshold. We again observe significant bunching at £300,000.

Finally, we apply the methods recently developed in Chetty *et al.* (2011), Kleven (2016) and Best and Kleven (2017) to estimate the size of bunching in England formally. This also allows us to detect shifting of properties across the price distribution.

To do so, we construct a counterfactual price distribution for newly built units using information on sales excluding the region around the bunching thresholds. Following Kleven (2016), we estimate this counterfactual distribution by calculating the number of new build transactions in 100 GBP bins and use these to estimate:

$$S_l = \sum_{q=0}^{3} \phi_q (p_l - 600000)^q + \sum_{r \in R} \rho_r 1\left\{\frac{p_j}{r} \in \mathbb{N}\right\} + \varepsilon_l \quad (6)$$

where l indexes price bins. The dependent variable S_l measures the number of new build transactions in bin l. The first two sums provide an estimate of the counterfactual price distribution. The first sum is a third-degree polynomial on the difference between the price at bin l and the cut-off of £600,000. The second sum estimates fixed-effects for round numbers, with $\mathbb{N}$ representing the set of natural numbers and $R = \{5000, 10000, 25000, 50000\}$ representing a set of round numbers. The term ε_l corresponds to an error term. We estimate this equation with data for new build transactions in England taking place between April of 2013 (the introduction of HtB in England) and December of 2019 excluding transactions with prices between £590,000 and £630,000. We then obtain differences between this estimated counterfactual distribution and the observed distribution of prices to estimate bunching effects induced by HtB.

The difference between the size of the spike just under the threshold and the gap just after the threshold can be used to estimate the size of the local effect of HtB on new building activity. This can be driven by changes in the types of properties sold after accounting for local shifting in prices induced by the policy.

Appendix Figure B12 illustrates the difference between the observed density of property transactions and our estimated counterfactual density around the £600k notch. The counterfactual distribution is obtained by estimating equation (6). We observe substantial bunching below the cut-off of £600,000 and a large hole in the distribution above the cut-off. Using our counterfactual price distribution, we estimate there are 3,123 more transactions for properties valued from £590,000 to £600,000 and 1,272 *less* transactions for properties valued from £600,000 to £630,000.[31] These estimates suggest that HtB leads to a significant shift in housing construction away from properties above the price threshold, towards properties below the threshold. We relate this to changes in the size of built units in the next section.

6.2. Size Effect

We can also use our Difference-in-Discontinuities design to estimate the effect of HtB on the size (total floor area in square meters) of newly built housing units. We employ data on the size of new build transactions close to the GLA boundary and the English/Welsh border respectively, removing properties with prices over £600,000 and modifying equation (3) by using the log of unit size as dependent variable instead of the log of house prices. Due to planning and construction lags, we allow for a delayed response of one year so that the post-HtB period starts from April 2014 at the

[31] These numbers amount, respectively, to 12.3% and 5% of all sales in the £550000-£650000 range.

English/Welsh border and from February 2017 at the GLA boundary. Appendix Tables A8 and A9 summarize the estimation results.

Table A8 displays largely negative, albeit, with one exception, statistically insignificant effects of the London HtB scheme on the size of newly built units. Together with the results for price bunching in England, this suggests that developers may have adjusted the characteristics of properties to meet the HtB conditions. Yet, the fact that most coefficients are not significant, warrants some caution with this interpretation. Table A9 reports estimates of the size effect at the English/Welsh border; we again overwhelmingly find negative but statistically insignificant effects.

7. Back-of-the-Envelope Calculation: Price Effect vs. Interest Subsidy

Our preferred empirical estimate from Table 5 indicates that the introduction of London's HtB led to a 7.6% increase in house prices inside the GLA relative to outside the GLA. The policy effect can operate via two main channels. First, as discussed in the theoretical section, the reduction in required down-payments can increase demand, leading to higher prices. Second, the government equity loan has a lower interest rate than that typically paid for mortgage loans and during the first five years the equity loan carries no interest. This interest rate subsidy could also result in higher demand, and an increase in prices. Because both effects are bundled together with the policy, it is not easy to disentangle them empirically. However, we do know the size of the interest rate subsidy in each period. We can combine this with prevailing interest rates on mortgages, discount rates and reasonable assumptions for the expected appreciation of house prices to obtain the present discounted value of that subsidy. Using these numbers, we can decompose the 7.6% total price effect into an interest-rate subsidy and a credit relaxation effect.

We compare discounted cash flows for two hypothetical households buying a property using different forms of financing. Household A buys a property using a London HtB equity loan for 40% of the property value, a 55% LTV mortgage and a 5% down-payment. Household B buys a property of the same price using an English HtB equity loan for 20% of the property value, a 75% LTV mortgage and a 5% down-payment.

For the sake of simplicity, we consider a 10-year time window.[32] We also assume that both mortgages are 10 year fixed-rate with equal rates. Under these assumptions, the value of the interest-rate subsidy accruing to household A can be obtained by comparing two figures: i) the discounted value of payments for a 20% reduced-interest HtB-equity loan after subtracting the proceeds from saving the 20% cash excess in a standard household portfolio, and ii) the stream of payments arising from a 20% non-HtB 10-year fixed rate mortgage.

HtB equity mortgages require no interest paid in the first 5 years since the purchase. After that, the interest rate is 1.75% × (1 + (1%+Retail Price Inflation)). Assuming the RPI stays constant at 3.1% (the average rate between 2010 and 2019), we can trace out future payments on all HtB loans. We assume a 10-year fixed-rate mortgage pays a nominal interest of 3.27% based on the Bank of England data of rates for January 2016. This pins down the path of interest payments for mortgages. The interest on savings is assumed to be equal to 1.4%, which is also taken as the discount rate in our present value calculations.[33] Finally, the expected yearly growth rate for house prices – necessary to value foregone capital gains on the 20% equity stake of the government – is assumed to be 1.5%, which was the average growth rate of real house prices in the UK between 2010 and 2019.

Based on these assumptions, we find that the net present value of the interest rate subsidy is 4.11% of the value of the purchased property. This figure is 54% of our total estimated price effect, implying that about half of the effect of London's HtB on prices operates via a relaxation of household credit conditions.

The calculated present discounted value of the interest rate subsidy depends on our assumptions regarding mortgage rates, inflation expectations, etc. A sensitivity analysis reveals that this figure is particularly responsive to the expected appreciation

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[32] We assume household A pays off half of the equity loan in a single installment in year 10 and household B pays off an equivalent amount of the mortgage in the same way. After these payments, both agents are left with a HtB loan amounting to 20% of the property's initial price and a mortgage amounting to 55% of the initial price (minus any amortization paid in the intervening years). Note that if household A repays the HtB equity loan earlier – which is plausible, given the evidence in Benetton *et al.* (2019) – this would reduce the present discounted value of the interest rate subsidy even further.

[33] The interest on savings is computed by observing assets held by recent buyers – buyers purchasing property in the previous two years - as recorded in the 2016 Wealth and Assets Survey. Most households in this group have their wealth in a combination of traditional savings accounts and ISAs (Individual Savings Accounts). One-year limited access ISAs are particularly popular, and they pay an interest of roughly 1.4%. Only about 1 in 10 recent buyers holds stocks or other risky assets. The average portfolio of a recent buyer (by size) – which is not quite representative of the most common portfolio – includes 10% invested in stocks. This increases total return on savings to roughly 2.6%.

rate of house prices and the return on savings. A higher appreciation rate of prices reduces the value of the subsidy. A higher return on savings has the opposite effect. Yet, given the other parameters of the exercise, either the expected capital gains should be unreasonably small (i.e. lower than 0.5%), or the returns on investment unreasonably large (i.e. as large as mortgage rates), for the interest subsidy to explain away our price effect. Hence, we conclude that there is indeed 'overcapitalization' of interest rate subsidies and we interpret this overcapitalization as a result of reduced down-payment requirements.

8. Conclusion

In 2013 the UK government announced the HtB scheme, which provides different forms of assistance to households aiming to buy a property as owner-occupiers. We exploit differences in the intensity of implementation of the policy's Equity Loan Scheme across two regional boundaries to estimate the effect of the policy on the price of newly built homes and on construction volumes. We estimate different effects depending on the boundary under consideration. In the case of the GLA, we find that the more generous London HtB program led to higher new build prices but had no discernible effect on construction volumes. Both of these effects are arguably contrary to the policy's objectives which are to improve affordability and promote new construction.

The estimated effects of the policy are more encouraging in the relatively supply-elastic markets around the English/Welsh border, with no significant effect on prices and a substantial and statistically significant effect on construction activity. Yet, the housing affordability crisis in the UK tends to be most severe in the supply inelastic markets of the South East and especially in the GLA.

Our findings suggest that HtB has stimulated housing construction in the 'wrong areas'; that is, it has stimulated construction in areas where planning constraints are less rigid and it is therefore comparably easy to build, not in areas where productivity and employment concentration are highest and new housing is most needed. This is consistent with observed patterns in the intensity of HtB-construction across England and Wales, illustrated in Appendix Figure B13: The policy has led to the construction of housing outside of the green belt areas of the most productive agglomerations in the UK (London, Oxford and Cambridge). This is in line with other stylized facts that suggest that workers increasingly commute excessively long distances through green belts to get from their place of residence to their workplace.

Contrary to the policy's title, HtB may not have 'helped' the population of credit constrained households in the most unaffordable areas of the country. There are two reasons for this. First, the policy pushed up house prices, increasing housing costs rather than housing consumption in square meters. Only developers or landowners, not new buyers, benefited from these policy-induced price increases. The price effect limits substantially the impact of the policy on the affordability conditions faced by credit constrained households. Second, the design of the HtB Equity Loan Scheme is such that those borrowers who took advantage of the scheme to gain access to the owner-occupied housing ladder, unlike existing homeowners, do not participate in the same way in future capital gains. This is because, at the time of sale, they have to pay back the equity loan at market value. If the price increases, so does the amount that the borrower owes the government. Ultimately, HtB arguably did little to 'help' young credit constrained households in unaffordable areas.

So, who benefited from HtB, if not the credit constrained households in the most unaffordable areas? Landowners in supply constrained areas (including developers who held land in those areas prior to the policy's implementation) are amongst main beneficiaries. Moreover, our analysis of the financial performance of developers indicates that developers benefited too. This suggests that HtB not only had limited effects on affordability but may have also led to unwanted regressive distributional effects.

Tables

Table 1: Equity Loan Scheme in Different Regions in UK (applies to new build only)

Region	Introduction date	House value up to	Loan from government
England	April 2013	£600,000	Up to 20%
London	February 2016	£600,000	Up to 40%
Wales	January 2014	£300,000	Up to 20%

Table 2: Descriptive Statistics (Regression Sample)

	Observations	Mean	SD	Max	Min
Panel A: London, price effect					
House price (£)	41357	389440	279818	7850000	27720
HtB treatment	41357	0.3	0.4	1	0
Inside GLA	41357	0.6	0.5	1	0
Post London HtB	41357	0.5	0.5	1	0
Total floor area (m^2)	41357	86.8	48.5	797.5	1.5
Terrace	41357	0.2	0.4	1	0
Flat	41357	0.7	0.5	1	0
Detached	41357	0.1	0.3	1	0
Semi-detached	41357	0.1	0.3	1	0
Leasehold	41357	0.7	0.5	1	0
Energy consumption (kWh/m^2)	41357	102.2	69.2	1038	-128
Fireplace	41357	0.2	0.4	1	0
CO2 emissions (tons/year)	41357	1.4	1.1	36.9	-2
Distance to boundary (m)	41357	2515.4	1395.7	4999.3	4.7
Panel B: English/Welsh border, price effect					
House price (£)	11574	232536	111155	1554500	16260
HtB treatment	11574	0.4	0.5	1	0
Inside GLA	11574	0.5	0.5	1	0
Post English HtB	11574	0.8	0.4	1	0
Total floor area (m^2)	11574	101.4	40.7	575	3.5
Terrace	11574	0.2	0.4	1	0
Flat	11574	0.1	0.3	1	0
Detached	11574	0.5	0.5	1	0
Semi-detached	11574	0.2	0.4	1	0
Leasehold	11574	0.3	0.4	1	0
Energy consumption (kWh/m^2)	11574	106.3	47.8	1076	-31
Fireplace	11574	0.2	0.4	1	0
CO2 emissions (tons/year)	11574	1.9	1.2	61	-0.5
Distance to boundary (m)	11574	5028.3	2828.4	9981.3	11.2
Panel C: London, construction effect (ward-level sample)					
Number of units constructed	49320	0.8	3.3	93	0
Any new build in ward, by month	49320	0.2	0.4	1	0
HtB Treatment	49320	0.2	0.4	1	0
Inside GLA	49320	0.5	0.5	1	0
Post London HtB	49320	0.3	0.5	1	0
Distance to boundary (m)	49320	2773.4	1631.8	9214	186.7
Panel D: English/Welsh border, construction effect (ward-level sample)					
Number of units constructed	24480	0.5	1.7	73	0
Any new build in ward, by month	24480	0.2	0.4	1	0
HtB treatment	24480	0.2	0.4	1	0
In Wales	24480	0.6	0.5	1	0
Post HtB in England	24480	0.6	0.5	1	0
Distance to boundary (m)	24480	5439.9	3119.6	14592.8	324.2

Table 3: Descriptive Statistics (Developer Sample)

	Observations	Mean	SD	Max	Min
Panel A: HtB dummy sample					
Revenue (million £)	535	539.5	924.9	4874.8	1.8
Gross profit (million £)	535	110.8	212.4	1204.7	0.6
Net profit before tax (million £)	535	64.4	162.5	1090.8	0.03
HtB dummy	535	0.9	0.3	1	0
Panel B: HtB intensity sample					
Revenue (million £)	223	1150	1172.8	4874.8	6.1
Gross profit (million £)	223	223.1	275.8	1204.7	0.7
Net profit before tax (million £)	223	143.1	229.7	1090.8	0.1
HtB intensity	223	0.2	0.2	0.7	0
Panel C: HtB completion sample					
Revenue (million £)	493	573	954.2	4874.8	1.8
Gross profit (million £)	493	117.5	219.4	1204.7	0.6
Net profit before tax (million £)	493	68.5	168.3	1090.8	0.03
HtB completion ratio	493	0.4	0.1	0.4	0.2

Table 4: Supply Constraints Measures and Implied Price-Earnings Elasticities

Region	English/Welsh border	GLA boundary	Greater Manchester boundary
Share of land in green belts	3.77%	66.5%	52.6%
Average refusal rate 1979-2008	27.2%	35.6%	25.1%
Average share of developed land	6.3%	27.6%	18.2%
Average elevation range	476.0	143.9	382.3
Implied price-earning elasticity (OLS)	0.252	0.403	0.284
Implied price-earning elasticity (IV)	0.127	0.205	0.164

Notes: The refusal rate, share developed land and elevation range are weighted by the surface area of the Local Planning Authority. Data on refusal rates, share developed land and elevation range come from Hilber and Vermeulen (2016). The green belt shape file comes from the Ministry of Housing, Communities and Local Government.

Table 5: Price Effect at GLA Boundary

Specifications	(1)	(2)	(3)	(4)	(5)
HtB [1)]	0.0534**	0.0589***	0.0506*	0.0796***	0.0764***
	(0.0235)	(0.0222)	(0.0306)	(0.0268)	(0.0275)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	No	No	No
Housing controls [2)]	No	Yes	Yes	Yes	Yes
Postcode FEs	No	No	Yes	Yes	Yes
Census variables by year [3)]	No	No	No	Yes	Yes
Distance by year	No	No	No	No	Yes
N	41357	41357	41357	41357	41357
R^2	0.6388	0.6555	0.9277	0.9283	0.9283

Notes: All columns control for the logarithm of total floor area. [1)] HtB is a dummy taking value 1 for transactions inside the GLA after February 2016, when London's HtB was first introduced. [2)] Housing controls include dwelling type, the tenure of properties (freehold vs. leasehold), whether the property has a fireplace, energy consumption and CO_2 emissions. [3)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors are clustered at the postcode level.

Table 6: Price Effect at English/Welsh Border

Specifications	(1)	(2)	(3)	(4)	(5)
HtB [1)]	0.0801**	0.0682*	0.0012	0.0053	0.0090
	(0.0355)	(0.0348)	(0.0268)	(0.0283)	(0.0273)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	No	No	No
Housing controls [2)]	No	Yes	Yes	Yes	Yes
Postcode FEs	No	No	Yes	Yes	Yes
Census variables by year [3)]	No	No	No	Yes	Yes
Distance by year	No	No	No	No	Yes
N	11574	11574	11574	11574	11574
R^2	0.6683	0.7028	0.9167	0.9171	0.9173

Notes: All columns control for the logarithm of total floor area. [1)] HtB is a dummy taking value 1 for transactions in England after April 2013, when the English version of HtB was introduced. [2)] Housing controls include dwelling type, the tenure of properties (freehold vs. leasehold), whether the property has a fireplace, energy consumption and CO_2 emissions. [3)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors are clustered at the postcode level.

Table 7: Construction Effect at GLA Boundary

Dependent Variable:	#New builds				Dummy			
Specifications	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HtB [1)]	0.0015	0.0015	-0.1312	-0.0964	0.0094	0.0094	-0.0004	0.0006
	(0.1688)	(0.1695)	(0.1930)	(0.1909)	(0.0229)	(0.0230)	(0.0237)	(0.0236)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	No	No	No	Yes	No	No	No
Ward fixed effects	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Census variables by year [2)]	No	No	Yes	Yes	No	No	Yes	Yes
Distance by year	No	No	No	Yes	No	No	No	Yes
N	49320	49320	49320	49320	49320	49320	49320	49320
R^2	0.0115	0.1545	0.1568	0.1579	0.0130	0.1781	0.1806	0.1812

Note: The dependent variable in columns 1 to 4 corresponds to the number of new builds in a ward. The dependent variable in columns 5 to 8 corresponds to a dummy taking value 1 if there was any sale of new build properties in that ward-month pair. [1)] HtB corresponds to a dummy taking value 1 for wards inside the GLA after January 2017 – which is one year after the implementation of London's HtB. [2)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors are clustered at the ward level.

Table 8: Construction Effect at English/Welsh Border

Dependent Variable:	#New builds				Dummy			
Specifications	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HtB [1)]	0.3988^{***}	0.3988^{***}	0.3771^{***}	0.3951^{***}	0.0876^{***}	0.0876^{***}	0.0783^{**}	0.0797^{**}
	(0.1438)	(0.1444)	(0.1311)	(0.1286)	(0.0321)	(0.0322)	(0.0308)	(0.0308)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	No	No	No	Yes	No	No	No
Ward fixed effects	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Census variables by year [2)]	No	No	Yes	Yes	No	No	Yes	Yes
Distance by year	No	No	No	Yes	No	No	No	Yes
N	24480	24480	24480	24480	24480	24480	24480	24480
R^2	0.0269	0.2302	0.2371	0.2387	0.0251	0.2253	0.2309	0.2318

Note: The dependent variable in columns 1 to 4 corresponds to the number of new builds in a ward. The dependent variable in columns 5 to 8 corresponds to a dummy taking value 1 if there was any sale of new build properties in that ward-month pair. [1)] HtB corresponds to a dummy taking value 1 for wards in England after March 2014 – which is one year after the implementation of the English version of HtB. [2)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors are clustered at the ward level.

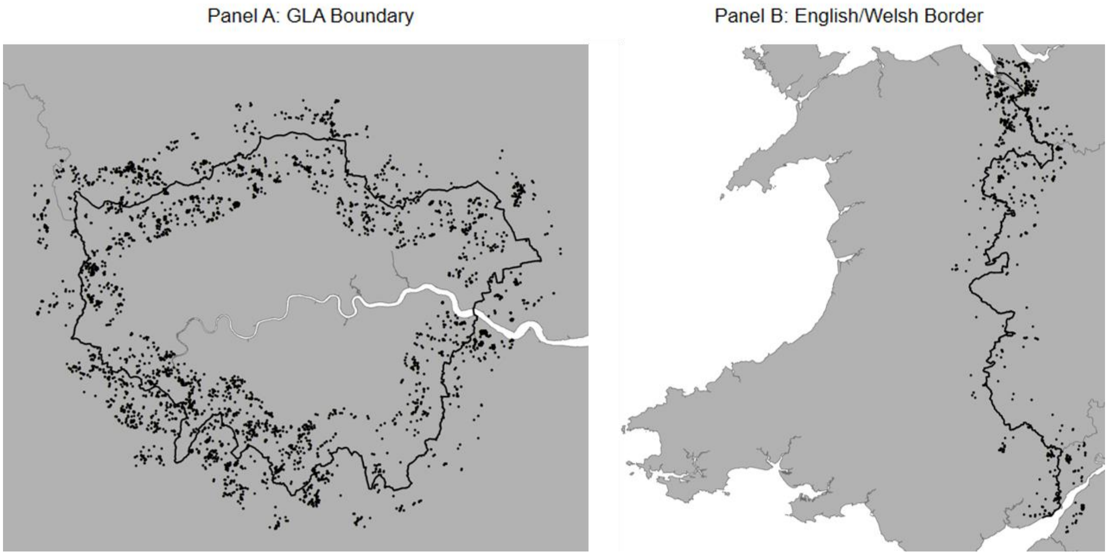
Table 9: Effects on Financial Performance of Developers

Specifications	(1)	(2)	(3)
Dependent variable	Ln(turnover)	Ln(gross profit)	Ln(net profit before tax)
	Panel A: DID sample (*N* = *535*)		
HtB dummy × Post [1)]	0.4509^{***}	0.7957^{***}	1.3686^{***}
	(0.1309)	(0.2711)	(0.4148)
	Panel B: HtB intensity sample (*N* = *223*)		
HtB intensity × Post [2)]	1.0086^{**}	1.6070^{**}	1.9312^{*}
	(0.4481)	(0.6659)	(1.1237)
	Panel C: Local HtB completion ratio as instrument for HtB dummy (*N* = *493*)		
HtB dummy × Post [1)]	0.6223^{**}	1.1813^{***}	1.4275^{**}
	(0.3113)	(0.4244)	(0.5767)
K.-P. F-statistics	4.842	4.842	4.842
Anderson-Robin 90% CI	(0.23,1.79)	(0.58,2.63)	(0.33,2.91)
	Panel D: first stage (*N* = *493*)		
HtB completion × Post [3)]		2.3997^{**}	
		(1.0906)	
	Panel E: Local HtB completion ratio (*N* = *493*)		
HtB completion × Post [3)]	1.4934^{**}	2.8349^{**}	3.4257^{*}
	(0.6763)	(1.2505)	(1.9808)
Year FEs	Yes	Yes	Yes
Developer FEs	Yes	Yes	Yes

Note: [1)] HtB dummy equals 1 if a developer is registered at one of the HtB regional offices. [2)] HtB intensity is defined as the 5-year average ratio of HtB-completions relative to all property completions by the developer. [3)] HtB completion ratio represents the number of HtB completions relative to the number of new build transactions at the NUTS-1 level. Standard errors are clustered at the developer level. K.-P. refers to Kleinbergen-Paap first-stage F-statistics.

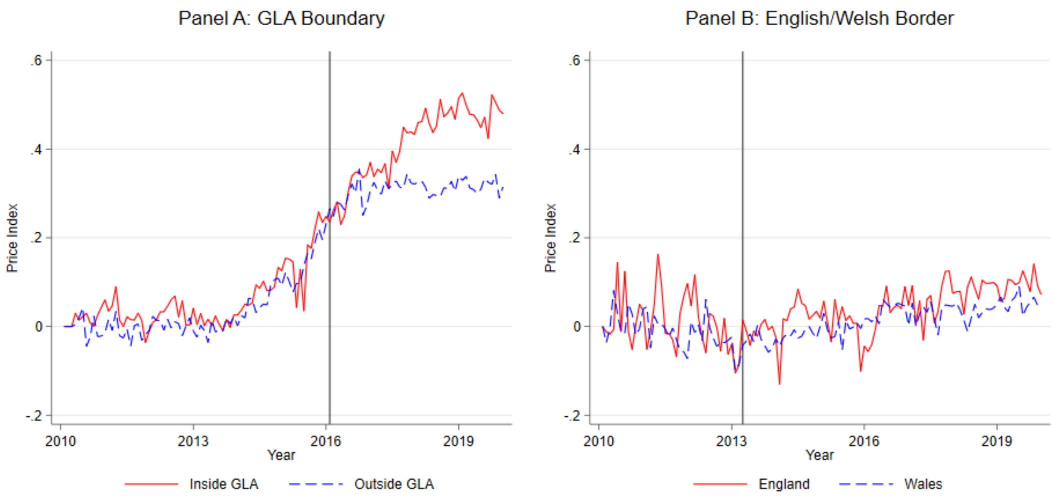
Figures

Fig. 1: New Builds near the GLA Boundary and English/Welsh Border



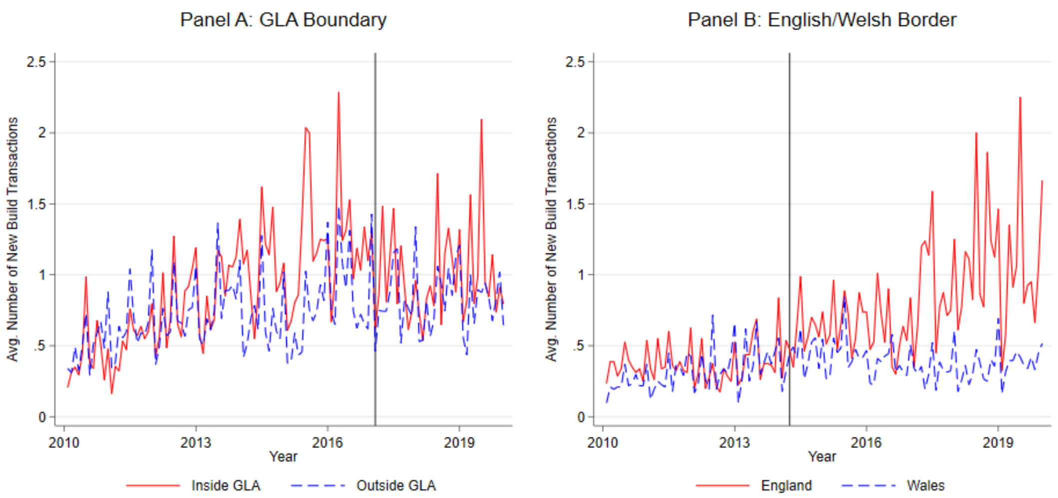
Note: In Panel A, solid black line represents the boundary of the Greater London Authority (GLA). Each of the black dots represents a new build sale taking place during our sample period within 5km of the boundary. In Panel B, solid black line represents the English-Welsh border. Each of the black dots represents a new build sale taking place during our sample period within 10km of the boundary.

Fig. 2: House Price Index



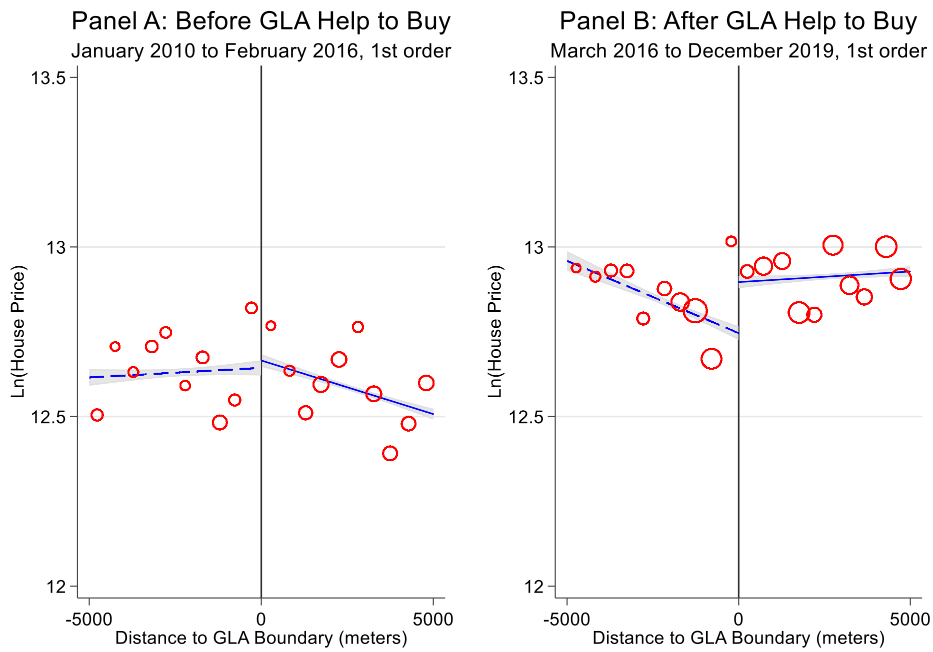
Note: In Panel A, the vertical line represents January 2016. In February 2016, London's Help to Buy scheme was implemented. In Panel B, the vertical line represents March 2013. In April 2013, Help to Buy was implemented in England.

Fig. 3: Housing Construction



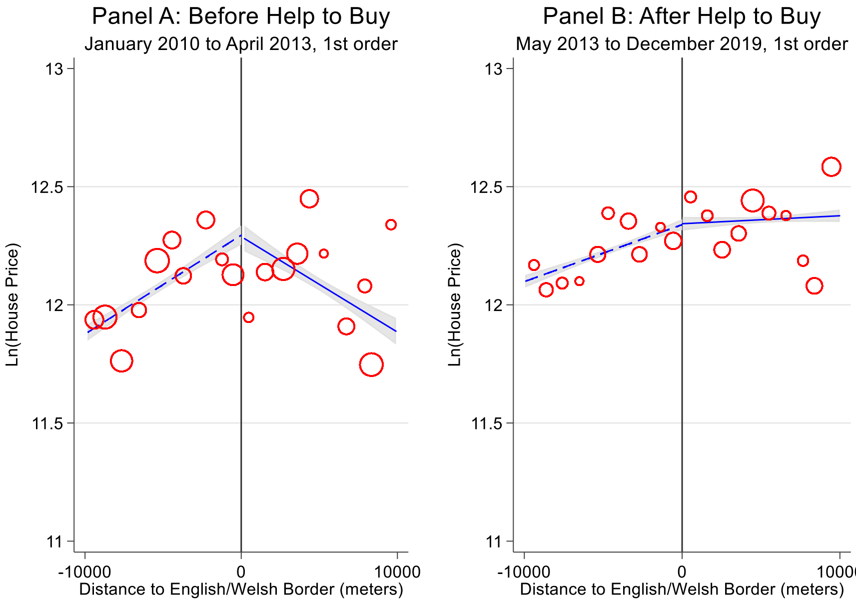
Note: In Panel A, the vertical line represents January 2017. In February 2016, London's Help to Buy scheme was implemented. In Panel B, the vertical line represents March 2014. In April 2013, Help to Buy was implemented in England.

Fig. 4: Price Effect at GLA Boundary (Boundary Discontinuity Design)



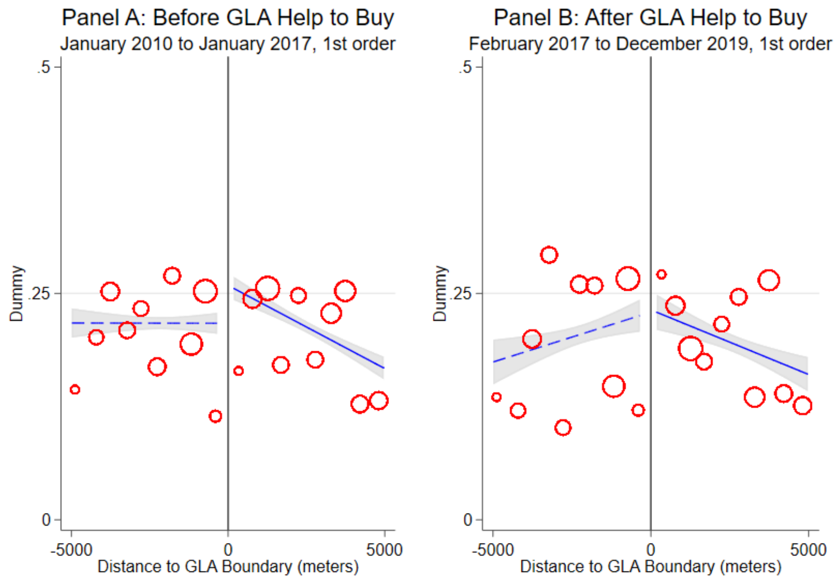
Note: Prices of new builds close to the GLA boundary. Positive distance: transactions inside the GLA; Negative distance: transactions outside the GLA. Circles represent averages taken within 0.5km bins, with the diameter of each circle corresponding to the number of sales in that bin. Lines correspond to first degree polynomial estimates separately on each side of the boundary. Shaded areas correspond to 95% confidence intervals around those lines.

Fig. 5: Price Effect at English/Welsh Border (Boundary Discontinuity Design)



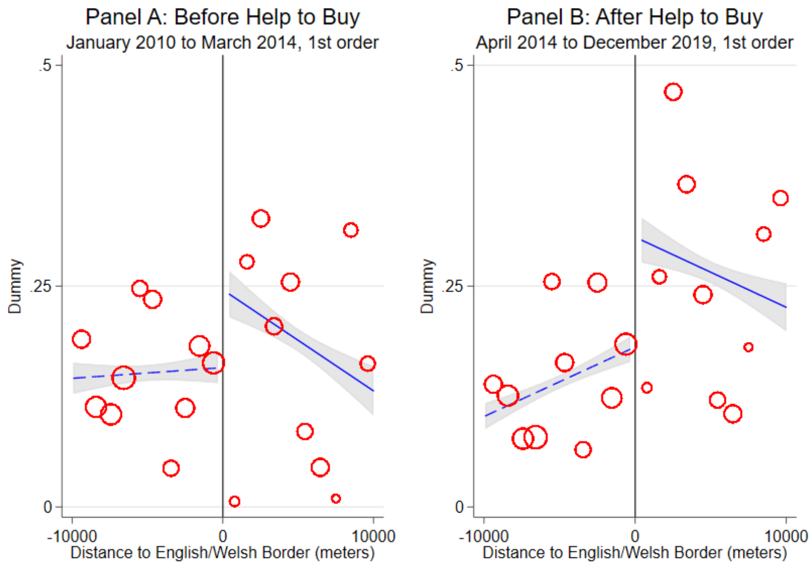
Note: Prices of new builds close to the English/Welsh border. Positive distance: transactions in England; Negative distance: transactions in Wales. Circles represent averages taken within 1km bins, with the diameter of each circle corresponding to the number of sales in that bin. Lines correspond to first degree polynomial estimates separately on each side of the boundary. Shaded areas correspond to 95% confidence intervals around those lines.

Fig. 6: Construction Effect at GLA Boundary (Boundary Discontinuity Design)



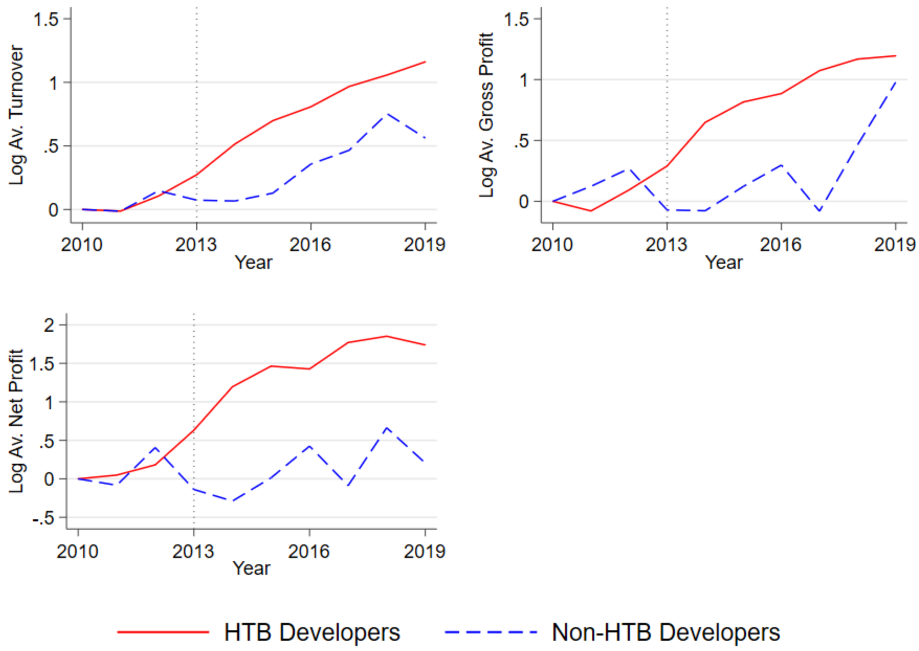
Note: Vertical axis corresponds to the average of a dummy taking value 1 for wards with at least one new build sale, 0 otherwise. Circles represent averages taken within 1km bins, with the diameter of each circle corresponding to the number of sales in that bin. Lines correspond to first degree polynomial estimates separately on each side of the boundary. Shaded areas correspond to 95% confidence intervals around those lines. Positive distance: wards inside the GLA; Negative distance: wards outside the GLA.

Fig. 7: Construction Effect at English/Welsh Border (Boundary Discontinuity Design)



Note: Vertical axis corresponds to the average of a dummy taking value 1 for wards with at least one new build sale, 0 otherwise. Circles represent averages taken within 1km bins, with the diameter of each circle corresponding to the number of sales in that bin. Lines correspond to first degree polynomial estimates separately on each side of the boundary. Shaded areas correspond to 95% confidence intervals around those lines. Positive distance: wards in England; Negative distance: wards in Wales.

Fig. 8: Developers' Financial Performance over Time



Note: Vertical axis represents the log of financial performance indicators computed after removing company fixed effects (normalized to 0 in 2010). The dotted line corresponds to data for all of 2013. Help to Buy was implemented in England in April 2013, hence the 2013 data contains both pre- and post-information.

Appendices

Appendix A: Appendix Tables

Table A1: Number of Transactions

London					
	5km	4 km	3 km	2 km	1 km
Total number of sales	41357	33031	24955	17736	6606
Postcodes	3751	3005	2276	1578	642
Sales in treatment group post London HtB	10812	7877	5982	3996	1602
Postcodes	744	540	415	299	122
Wales					
	10km	9km	8km	7km	6km
Total number of sales	11574	10362	8895	8085	7527
Postcodes	1101	974	868	789	713
Sales in treatment group post English HtB	5081	4283	3687	3430	3225
Postcodes in treatment group	492	414	368	348	316

Notes: Number of new build property sales for bands around the GLA (top panel) or Welsh (bottom panel) boundaries. Bandwidths in each case indicated in the top row of the bottom and top panels.

Table A2: Testing for Displacement Effect around GLA (Price Effect)

Sample:	Inside London		Outside London	
Specifications	(1)	(2)	(3)	(4)
Close x Post [1)]	-0.0237		0.0011	
	(0.0337)		(0.0350)	
Distance x Post [2)]		0.0185*		-0.0091
		(0.0100)		(0.0128)
Year-month fixed effects	Yes	Yes	Yes	Yes
Housing Controls [3)]	Yes	Yes	Yes	Yes
Postcode FEs	Yes	Yes	Yes	Yes
Census variables by year [4)]	Yes	Yes	Yes	Yes
N	24274	24274	17083	17083
R^2	0.9134	0.9134	0.9500	0.9500

Notes: Sample in columns 1 and 2 corresponds to properties sold inside of London within 5km of the GLA boundary. Sample in columns 3 and 4 corresponds to properties sold outside of London within 5km of the GLA boundary. Dependent variable is the logarithm of sale price in all columns. [1)] Close is a dummy taking value 1 for properties within 2.5km of the GLA boundary. Post is a dummy that takes value 1 if an individual transaction occurs after February 2016. [2)] Distance represents the distance from a property to the GLA boundary. [3)] Housing controls include the logarithm of total floor area, dwelling type, the tenure of properties (freehold vs. leasehold), whether the property has a fireplace, energy consumption and CO_2 emissions. [4)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors clustered at the postcode level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A3: Displacement Effect around English/Welsh Border (Construction Effect)

Sample:	England		Wales	
Specifications	(1)	(2)	(3)	(4)
Close x $Post_{t-12}$ [1)]	0.0153		0.0378	
	(0.0531)		(0.0415)	
Distance x $Post_{t-12}$ [2)]		-0.0014		-0.0063
		(0.0076)		(0.0068)
Year-month fixed effects	Yes	Yes	Yes	Yes
Ward FEs	Yes	Yes	Yes	Yes
Census variables by year [3)]	Yes	Yes	Yes	Yes
N	9600	9600	14880	14880
R^2	0.2379	0.2379	0.2281	0.2281

Notes: Sample in columns 1 and 2 corresponds to wards in England within 10km of the English/Welsh border. Sample in columns 3 and 4 corresponds to wards in Wales within 10km of the English/Welsh border. Distance between ward and border calculated from the ward's centroid. Dependent variable is a dummy variable that takes the value 1 if there was any sale of new build properties in that ward-month pair. [1)] Close is a dummy taking value 1 for wards with centroids within 5km of the English/Welsh border. Post is a dummy that takes value 1 for year-month fixed effects post March 2014 (one year after the implementation of the English HtB). [2)] Distance represents the straight-line distance between the ward centroid and the English/Welsh border. [3)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors clustered at the ward level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A4: Placebo - Price Effect at Greater Manchester Boundary

Specifications	(1)	(2)	(3)	(4)	(5)
HtB (placebo) [1)]	0.0306	0.0186	-0.0151	-0.0025	-0.0037
	(0.0328)	(0.0320)	(0.0208)	(0.0234)	(0.0228)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	No	No	No
Housing controls [2)]	No	Yes	Yes	Yes	Yes
Postcode fixed effects	No	No	Yes	Yes	Yes
Census variables by year [3)]	No	No	No	Yes	Yes
Distance by year	No	No	No	No	Yes
N	18142	18142	18142	18142	18142
R^2	0.5984	0.6532	0.9157	0.9160	0.9161

Notes: Estimates obtained on the sample of new build properties within 5km of the Greater Manchester boundary between 2010 and 2019. Dependent variable is the logarithm of the transaction price. All columns control for the logarithm of total floor area. [1)] HtB variable takes value 1 for properties sold in Greater Manchester after February 2016. [2) - 3)] see Table 5. Standard errors clustered at the postcode level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A5: Placebo - Construction Effect at Greater Manchester Boundary

Dependent Variable:	#New builds				Dummy			
Specifications	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HtB [1]	0.2711	0.2711	0.3674	0.3534	-0.0227	-0.0227	-0.0031	-0.00001
	(0.2602)	(0.2613)	(0.2892)	(0.2861)	(0.0433)	(0.0435)	(0.0448)	(0.0443)
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	No	No	No	Yes	No	No	No
Ward FEs	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Census variables by year [2]	No	No	Yes	Yes	No	No	Yes	Yes
Distance by year	No	No	No	Yes	No	No	No	Yes
N	21480	21480	21480	21480	21480	21480	21480	21480
R^2	0.0422	0.2242	0.2283	0.2292	0.0355	0.2632	0.2685	0.2694

Note: Estimated over the sample of wards within 5km of the Greater Manchester boundary between 2010 and 2019. The dependent variable in columns 1 to 4 corresponds to the number of new builds in a ward. The dependent variable in columns 5 to 8 corresponds to a dummy taking value 1 if there was any sale of new build properties in that ward-year pair. [1] HtB corresponds to a variable taking value 1 for wards inside of Greater Manchester after January 2017 (1 year after the introduction of London's HtB). [2] See Table 7. Standard errors clustered at the ward level in parentheses.

Table A6: Additional Robustness Checks – GLA Boundary

Dependent Variable:	Ln(house price)		#New builds	Dummy
Specifications	(1)	(2)	(3)	(4)
	2.5 km bandwidth	7.5 km bandwidth	Contemporaneous effect	Contemporaneous effect
HtB [1)]	0.0832*	0.0520**	0.0106	0.0068
	(0.0429)	(0.0249)	(0.1726)	(0.0234)
Year-month FEs	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	Yes	Yes
Housing controls [2)]	Yes	Yes	No	No
Postcode FEs	Yes	Yes	No	No
Ward FEs	No	No	Yes	Yes
Census variables by year [3)]	Yes	Yes	Yes	Yes
Distance by year	Yes	Yes	Yes	Yes
N	21456	65140	49320	49320
R^2	0.9341	0.9205	0.1578	0.1812

Note: Columns 1 and 2 control for the logarithm of total floor area. The dependent variable in column 3 corresponds to the number of new builds in a ward-month pair, and the dependent variable in column 4 corresponds to a dummy taking value 1 if there was any sale of new build properties in that ward-year pair. [1)] In columns 1 and 2, HtB is a dummy taking value 1 for transactions inside of GLA after February 2016. In columns 3 and 4, HtB is a dummy taking value 1 for wards inside the GLA after January 2017. [2) - 3)] See Table 5. Standard errors clustered at either the postcode level (in columns 1 and 2) or the ward level (in columns 3 and 4) in parentheses. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A7: Additional Robustness Checks – English/Welsh Border

Dependent Variable:	Ln(house price)			#New builds		Dummy	
Specifications	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	5 km bandwidth	15 km bandwidth	Restricted sample [4]	Contemp. effect	Restricted sample	Contemp. effect	Restricted sample
HtB [1]	0.0144	-0.0103	0.0242	0.3364***	0.4162***	0.0675**	0.0820***
	(0.0340)	(0.0206)	(0.0360)	(0.1281)	(0.1392)	(0.0292)	(0.0314)
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Housing controls [2]	Yes	Yes	Yes	No	No	No	No
Postcode FEs	Yes	Yes	Yes	No	No	No	No
Ward FEs	No	No	No	Yes	Yes	Yes	Yes
Census variables by year [3]	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance by year	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	6310	19302	10763	24480	22644	24480	22644
R^2	0.9093	0.9027	0.9176	0.2376	0.2377	0.2309	0.2304

Note: Columns 1, 2, and 3 control for the logarithm of total floor area. The dependent variable in columns 4 and 5 corresponds to the number of new builds in a ward-month, and the dependent variable in columns 6 and 7 corresponds to a dummy taking value 1 if there was any sale of new build properties in that ward-year pair. [1] HtB is a dummy taking value 1 either for transactions in England after April 2013 (columns 1 to 3), or for wards in England after March 2014 (columns 4 to 7). [2) - 3)] See Table 5. [4] The restricted sample excludes new build transactions between April 2013 and December 2013, when HtB was only introduced in England but not in Wales. Standard errors clustered at either the postcode level (columns 1 to 3) or the ward level (columns 4 to 7) in parentheses. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A8: Size Effect at GLA Boundary

Specifications	(1)	(2)	(3)	(4)	(5)
HtB [1)]	0.0224	-0.0024	-0.0242	-0.0303*	-0.0298
	(0.0266)	(0.0131)	(0.0183)	(0.0172)	(0.0188)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	No	No	No
Housing controls [2)]	No	Yes	Yes	Yes	Yes
Postcode FEs	No	No	Yes	Yes	Yes
Census variables by year [3)]	No	No	No	Yes	Yes
Distance by year	No	No	No	No	Yes
N	36930	36930	36930	36930	36930
R^2	0.4198	0.7770	0.9344	0.9346	0.9347

Notes: Sample corresponds to new build property sales under £600,000 within 5km of the GLA boundary for the period between 2010 and 2019. Dependent variable is the logarithm of total floor area in square meters. All columns control for the logarithm of transaction price. [1)] HtB corresponds to a dummy taking value 1 for properties sold inside of London after January 2017 – which is one year after the implementation of London's HtB. [2)] Housing controls include dwelling type, the tenure of properties (freehold vs. leasehold), whether the property has a fireplace, energy consumption and CO_2 emissions. [3)] Neighbourhood controls (from the 2011 Census) are the percentage of (1) married residents and (2) residents with level-4 and above educational qualifications at ward level. Standard errors are clustered at the postcode level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

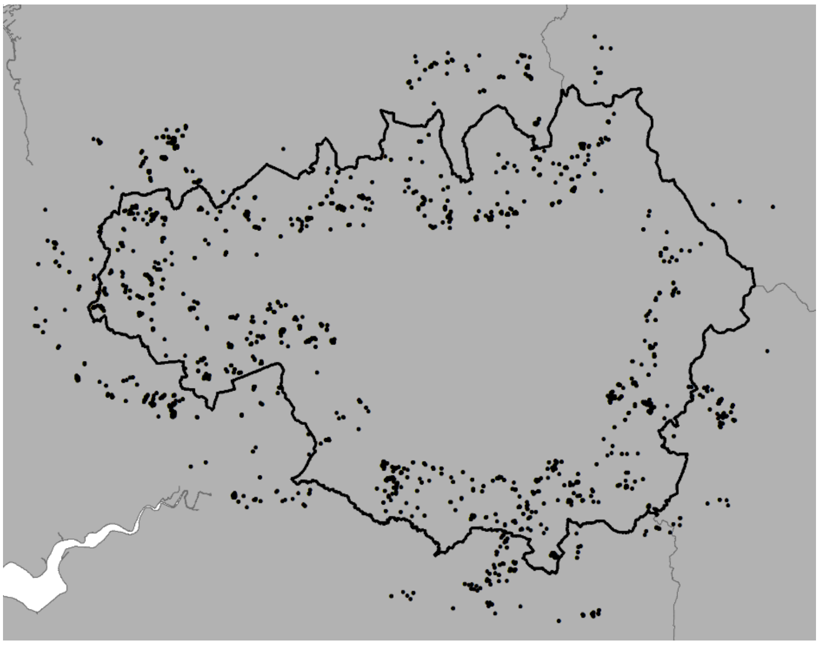
Table A9: Size Effect at English/Welsh Border

Specifications	(1)	(2)	(3)	(4)	(5)
HtB [1)]	-0.0398	-0.0392**	-0.0152	-0.0241	-0.0258
	(0.0276)	(0.0192)	(0.0241)	(0.0239)	(0.0228)
Year-month fixed effects	Yes	Yes	Yes	Yes	Yes
Distance to boundary on each side	Yes	Yes	No	No	No
Housing controls [2)]	No	Yes	Yes	Yes	Yes
Postcode FEs	No	No	Yes	Yes	Yes
Census variables by year [3)]	No	No	No	Yes	Yes
Distance by year	No	No	No	No	Yes
N	11490	11490	11490	11490	11490
R^2	0.6374	0.7873	0.9100	0.9105	0.9107

Notes: Sample corresponds to new build transactions under £600,000 within 10km of the English/Welsh border from 2010 to 2019. Dependent variable is the logarithm of total floor area in square meters. All columns control for the logarithm of transaction price. [1)] HtB is a dummy taking value 1 on the English side of the boundary after March 2014, which is one year after the implementation of the English version of HtB. [2) - 3)] See Table A8. Standard errors are clustered at the postcode level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

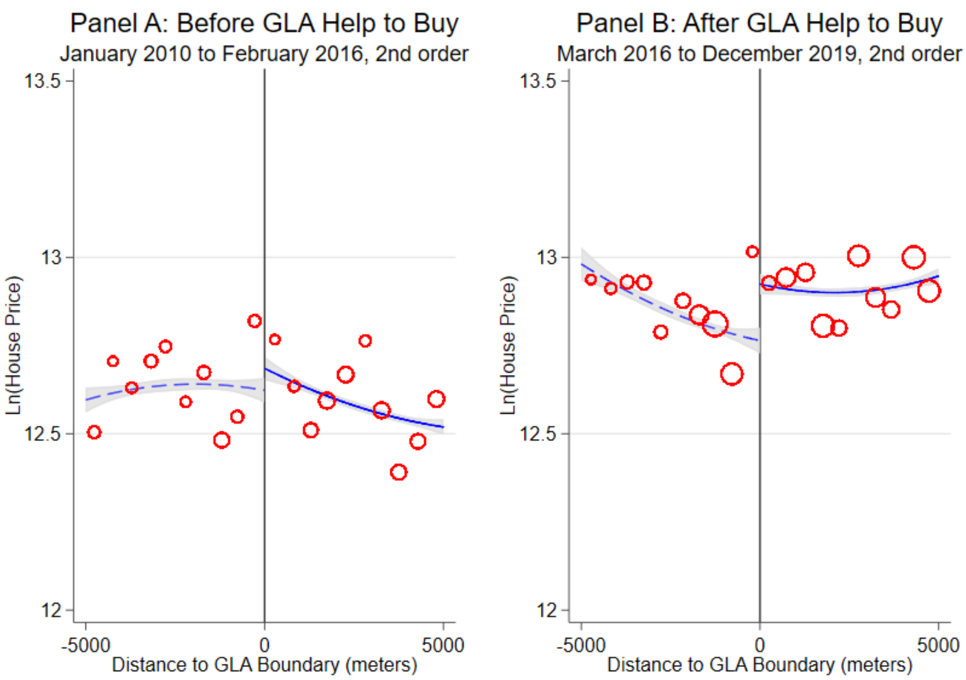
Appendix B: Appendix Figures

Fig. B1: New Builds near Greater Manchester Boundary



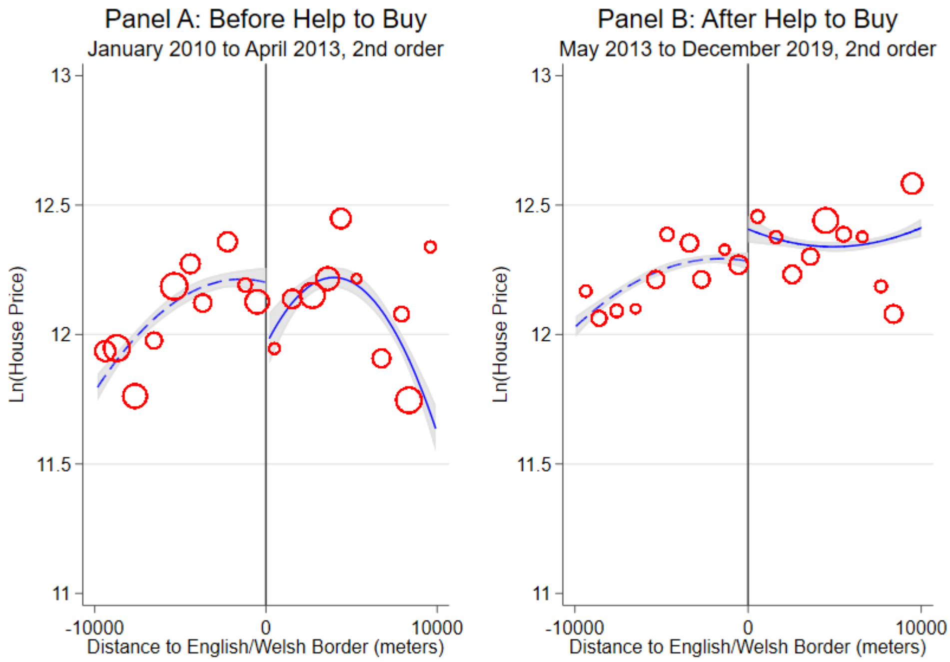
Note: Solid black line represents the Greater Manchester boundary. Each of the black dots represents a new build sale taking place during our sample period within 5km of the boundary.

Fig. B2: BDD Robustness – GLA Boundary HtB Price Effect, Second Order Polynomial



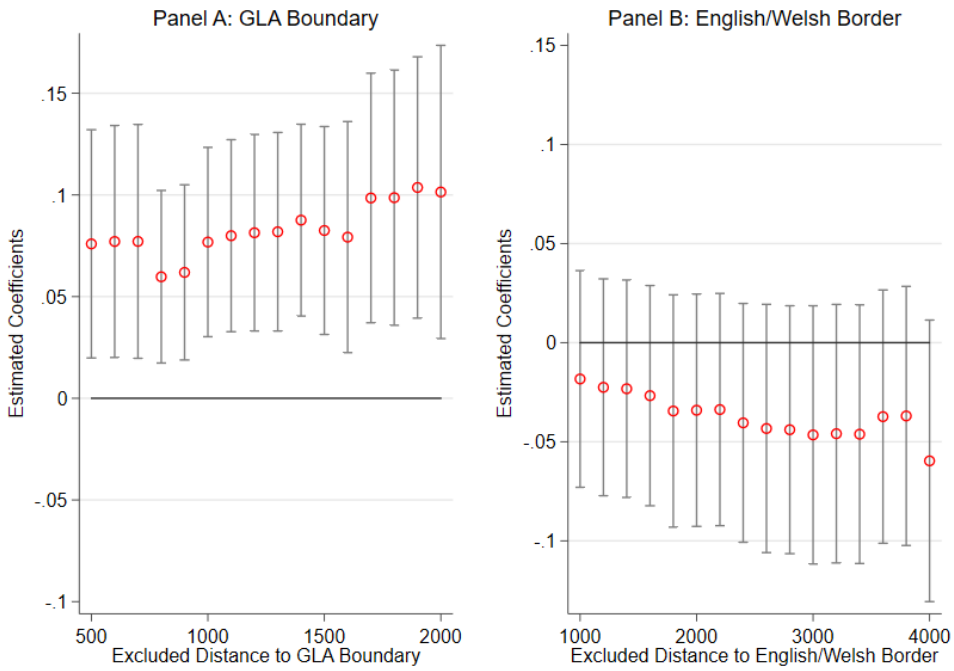
Note: Positive distance: transactions inside the GLA; Negative distance: transactions outside the GLA.

Fig. B3: BDD Robustness – English/Welsh Border Price Effect, Second Order Polynomial



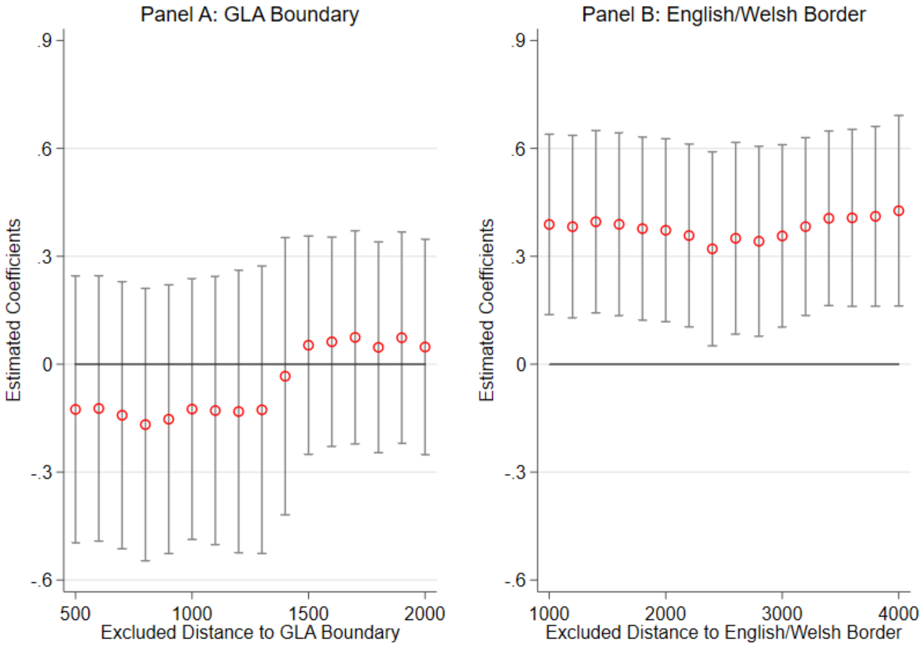
Note: Positive distance: transactions in England; Negative distance: transactions in Wales.

Fig. B4: Price Effect Excluding Properties near the Boundary



Note: Properties near the GLA boundary and the English/Welsh border are dropped respectively for the re-estimation of HtB's price effect. The horizontal axis represents the excluded distance. Red points correspond to the estimates of the price effect. Vertical lines correspond to 95% confidence intervals around those estimates.

Fig. B5: Construction Effects Excluding Wards near the Boundary

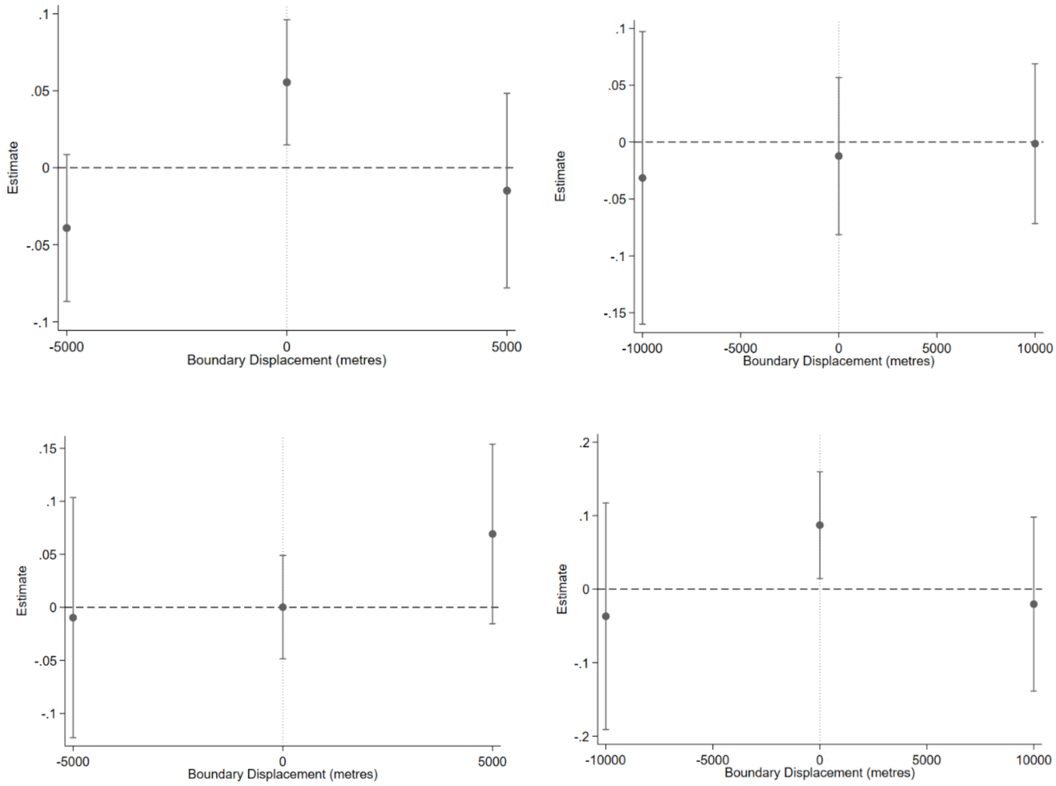


Note: Properties near the GLA boundary and the English/Welsh border are dropped respectively for the re-estimation of HtB's construction effect. The horizontal axis represents the excluded distance. Red points correspond to the estimates of the construction effect. Vertical lines correspond to 95% confidence intervals around those estimates.

Fig. B6: Boundary Placebos for England and Wales

Panel A: Price Effect in London

Panel B: Price Effect in Wales



Panel C: Construction Effect in London

Panel D: Construction Effect in Wales

Note: Main estimates and placebos for the price and construction estimates in England and Wales. Dots correspond to estimates under the full set of controls for each specification. Vertical bars correspond to 95% confidence intervals. Horizontal axes represent the displacement of the corresponding boundary relative to its original position. Estimated values at 0 correspond to the estimates reported in the right-most column of Tables 5 to 8. Negative values of the horizontal axis correspond to displacement into the side of the boundary with the less generous HtB scheme. Positive values correspond to displacement of the border towards the side of the boundary with the more generous HtB scheme.

Fig. B7: Histogram of House Prices in England

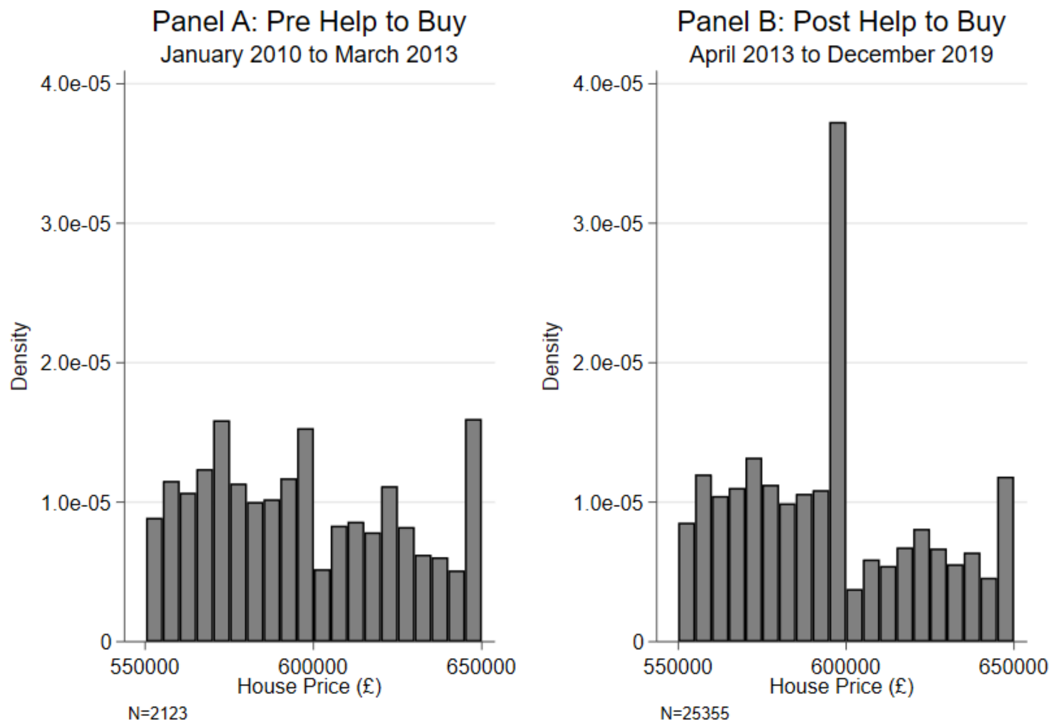


Fig. B8: Histogram of House Prices in Wales

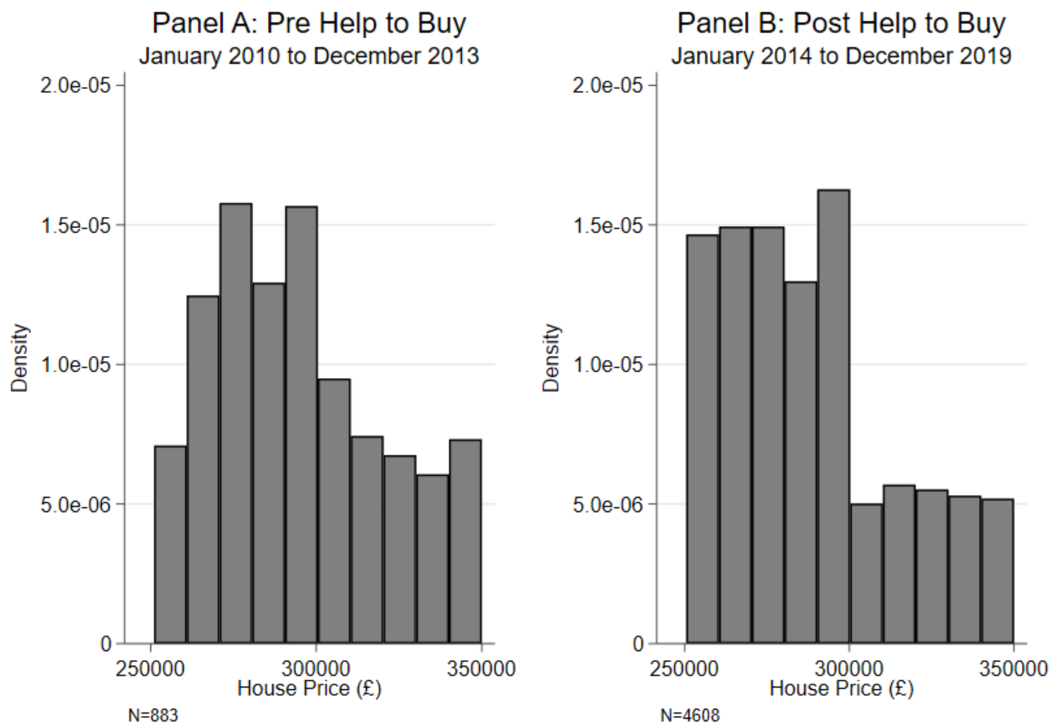
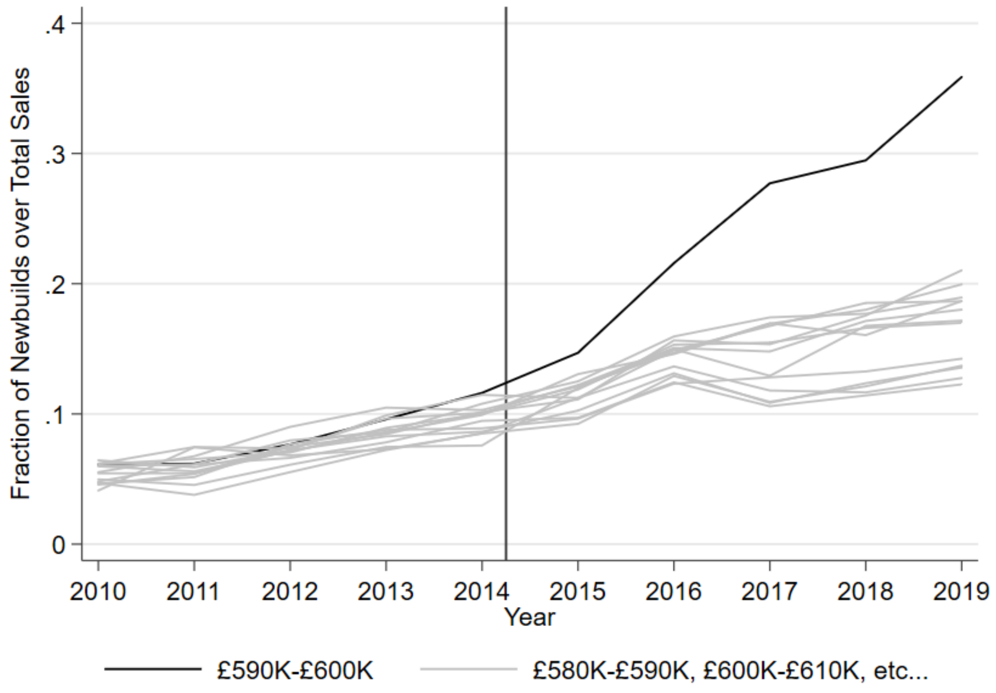
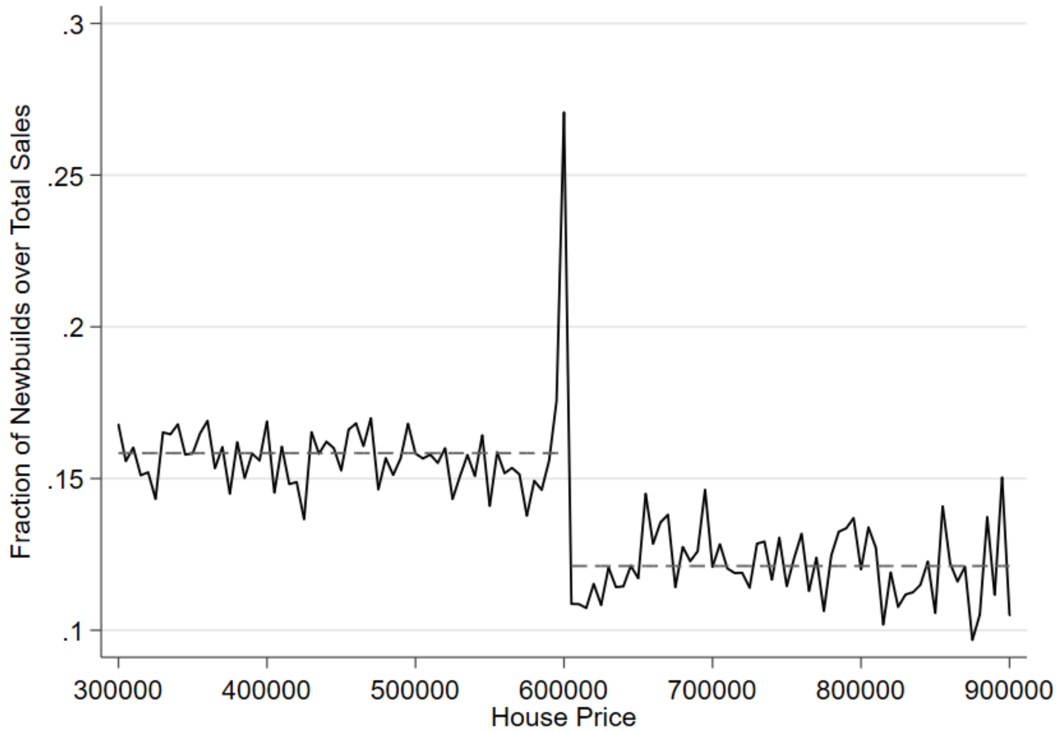


Fig. B9: Fraction of New Builds over Total Sales in England



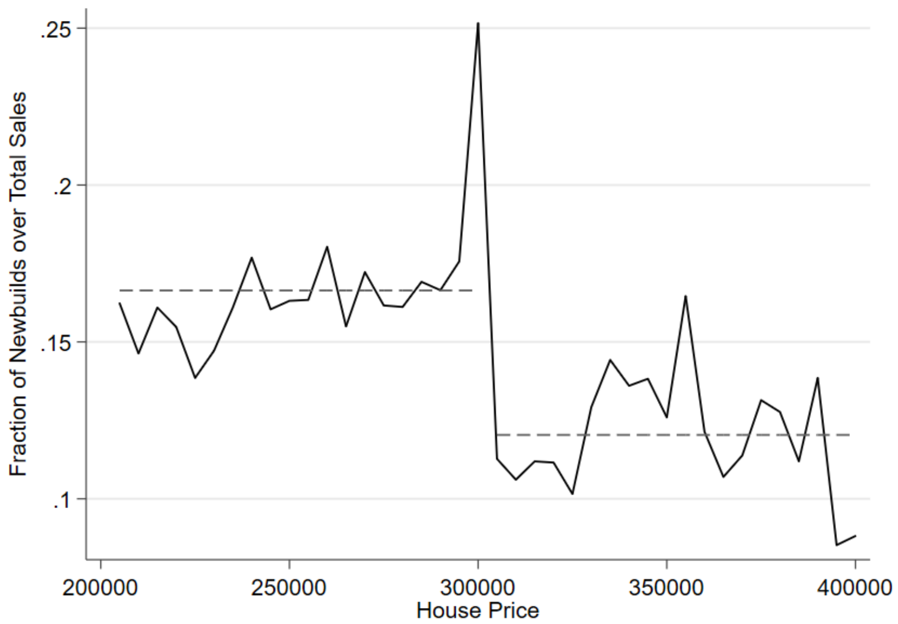
Note: The vertical line represents March 2014. In April 2013, Help to Buy was implemented in England.

Fig. B10: The Fraction of New Builds over Total Sales in England



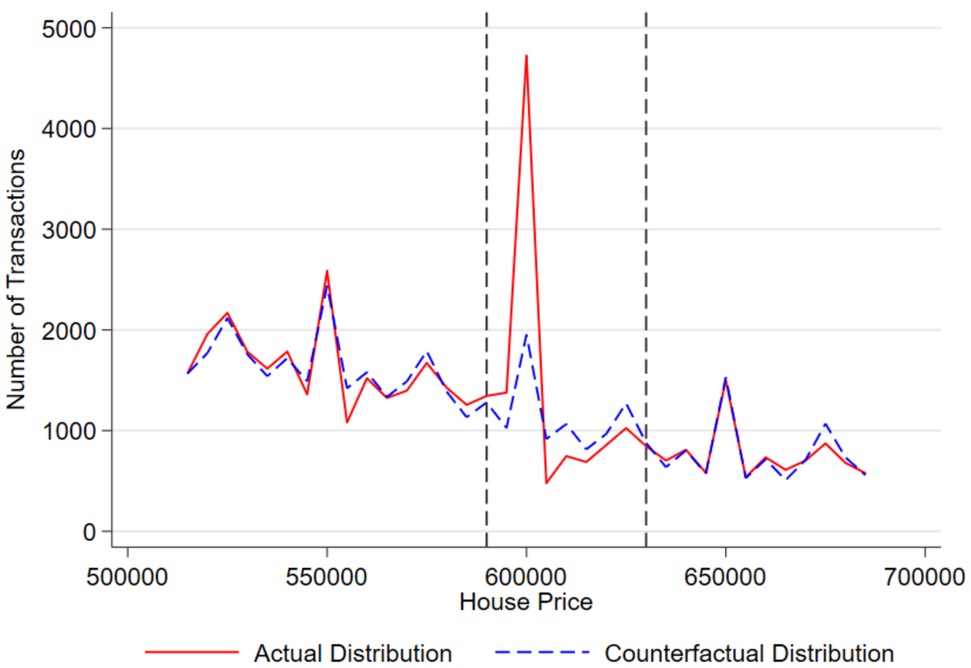
Note: Vertical axis measures fraction of new build sales over total sales for £5000 width price bins. Sales counted in the period between April 2013 and December 2019. England only.

Fig. B11: The Fraction of New Builds over Total Sales in Wales



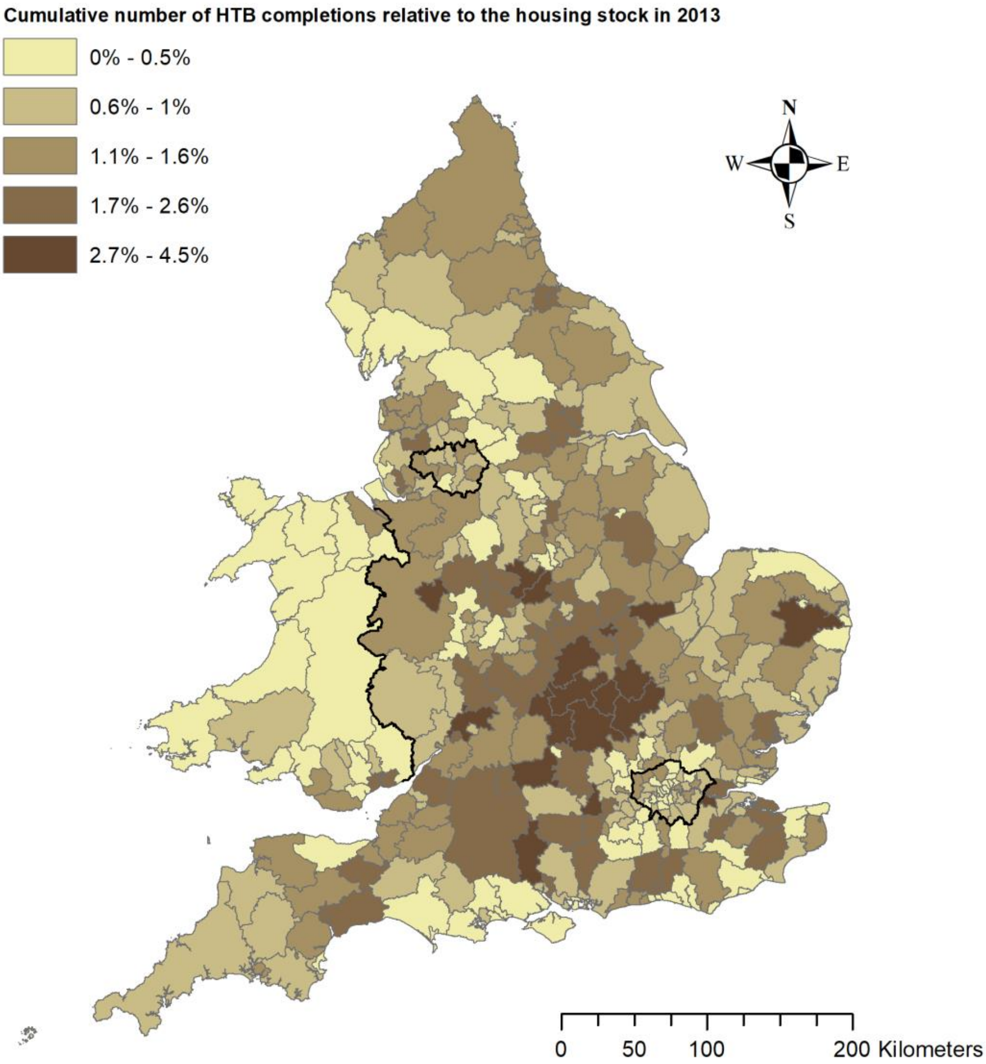
Note: Vertical axis measures fraction of new build sales over total sales for £5000 width price bins. Sales counted in the period between January 2014 and December 2019. Wales only.

Fig. B12: Estimated Bunching Effect



Note: Counter-factual distribution of prices estimated after excluding transactions between £590k and £630k, and represented using a dashed line. Other details on estimation discussed in the text.

Fig. B13: Accumulated Help to Buy Completions (2014-2019)



Appendix C: Theoretical Appendix

Demand for New Build Housing

Households buy a new build unit if:

$$u(c_{1h}, h) + \beta u(c_{2h}, h) \geq u(c_{1\emptyset}, 0) + \beta u(c_{2\emptyset}, 0)$$

$$w \geq (P - e)(1 + r)$$

$$e \geq P(1 - \gamma)$$

The first condition simply states that buying a new build is incentive compatible. Choice variables c_{th} and $c_{t\emptyset}$, correspond to consumption in period $t = \{1,2\}$ for households buying a new build and households renting, respectively. For a sufficiently large value of h, this condition is always satisfied given the assumption in footnote 9. The second condition is required to ensure households buying property in period 1 are able to meet their liabilities in period 2. This is satisfied given that $P < \frac{1}{1-\gamma}$ and assumption $w > \frac{\gamma}{1-\gamma}(1 + r)$. Finally, the third condition determines demand $Q_D = 1 - F_e((1 - \gamma)P)$.

Derivatives of Equilibrium Price and Quantities w.r.t. γ and ν

Competitive equilibrium results in $Q^* = \left(1 - F_e\big((1 - \gamma)\nu Q^*\big)\right)$. Total differentiation w.r.t. γ leads to:

$$\frac{dQ^*}{d\gamma} + f_e\big((1 - \gamma)\nu Q^*\big)\nu \left[(1 - \gamma)\frac{dQ^*}{d\gamma} - Q^*\right] = 0 \qquad \text{(A.1)}$$

$$\Rightarrow \frac{dQ^*}{d\gamma} = \frac{\nu f_e\big((1-\gamma)\nu Q^*\big)Q^*}{1 + f_e\big((1-\gamma)\nu Q^*\big)\nu(1-\gamma)} > 0$$

Similarly, total differentiating the equilibrium equation w.r.t. ν and re-arranging terms we obtain:

$$\frac{dQ^*}{d\nu} = \frac{-(1 - \gamma)f_e\big((1 - \gamma)\nu Q^*\big)Q^*}{1 + f_e\big((1 - \gamma)\nu Q^*\big)\nu(1 - \gamma)} < 0$$

To obtain the derivatives for equilibrium prices, note that the supply schedule is $P = \nu Q$. Therefore, we will have that:

$$\frac{dP^*}{d\gamma} = \nu \frac{dQ^*}{d\gamma} > 0 \quad \text{and} \quad \frac{dP^*}{d\nu} = Q^* + \nu \frac{dQ^*}{d\nu} = \frac{1}{1+f_e\big((1-\gamma)\nu Q^*\big)\nu(1-\gamma)} > 0$$

Proof of Proposition 1

Consider the general case in which e is distributed according to a general probability density function $f_e(\cdot)$. If we differentiate (A.1) with respect to ν, we obtain:

$$\big[1 + \nu(1-\gamma) f_e\big((1-\gamma)\nu Q^*\big)\big] \frac{dQ^*}{d\gamma d\nu} =$$

$$f_e\big((1-\gamma)\nu Q^*\big) \left[Q^* + \nu \frac{dQ^*}{d\gamma} - (1-\gamma)\frac{dQ^*}{d\nu}\right]$$

$$-f_e'\big((1-\gamma)\nu Q^*\big)\nu \left[(1-\gamma)Q^* + \nu(1-\gamma)\frac{dQ^*}{d\nu}\right]\left[(1-\gamma)\frac{dQ^*}{d\gamma} - Q^*\right] \qquad \text{(A.2)}$$

The term in square brackets in the first line is strictly positive. The term in square brackets in the second line is strictly *negative* as long as $\nu(1-\gamma) > f_e(\cdot)\ \forall e$. To see this, simply replace the expressions for $\frac{dQ^*}{d\gamma}$ and $\frac{dQ^*}{d\nu}$ above. Finally, if this condition holds, the first term in square brackets in the first line is positive and the second is negative.

Note that, under a uniform distribution of e, the term $f_e'\big((1-\gamma)\nu Q^*\big)$ is equal to 0. Therefore $\frac{dQ^*}{d\gamma d\nu} < 0$. To prove that $\frac{dP^*}{d\gamma d\nu} > 0$, we differentiate the expression for the supply schedule $P = \nu Q$ by γ and ν to obtain:

$$\frac{dP^*}{d\gamma d\nu} = \frac{dQ^*}{d\gamma} + \nu \frac{dQ^*}{d\gamma d\nu} \qquad \text{(A.3)}$$

Replacing the expressions for $\frac{dQ^*}{d\gamma}$ and $\frac{dQ^*}{d\gamma d\nu}$ in the uniform case, we obtain $\frac{dP^*}{d\gamma d\nu} > 0$.

For a general pdf $f_e(\cdot)$ – as long as we assume $\nu(1-\gamma) > f_e(\cdot)\ \forall e$, we can operate with (A.2) and (A.3) to prove the following: If e is distributed according to pdf $f_e(\cdot)$ with $f_e(\cdot)$ strictly decreasing in e and $\nu(1-\gamma) > f_e^{-1}(e)\ \forall e$, then $\frac{dQ^*}{d\gamma d\nu} < 0$. If $f_e(\cdot)$ is strictly increasing in e and $\nu(1-\gamma) > f_e^{-1}(e)\ \forall e$, then $\frac{dP^*}{d\gamma d\nu} > 0$.

It is important to note that conditions imposed on $f_e(\cdot)$ are sufficient and not necessary. Consider the case in which $e \sim \beta_{3,1}$. This pdf is strictly *increasing*, yet it can be shown that an increase in γ from 0.6 to 0.7 will result in a smaller increase in quantities when $\nu = 4.5$ rather than $\nu = 4$.[34]

Proof of Proposition 2

If F_c is the distribution of costs, then total profits are given by $\Pi(\mathrm{P}) = \int_0^P P - c \, dF_c$. Given the assumption on F_c above, this boils down to $\Pi(\mathrm{P}) = \int_0^P P - \frac{c}{\nu} \, dc = \frac{2\nu-1}{2\nu} P^2$, which is strictly increasing in P. Given that $\frac{dP^*}{d\gamma} > 0$, it follows that $\frac{d\Pi(\mathrm{P}^*)}{d\gamma} > 0$.

[34] Code available upon request.

Chapter 2

Low-rise Buildings in Big Cities:

Theory and Evidence from China

1. Introduction

Land use regulation has been a long-time focus of economic research. It is imposed in every country in the world with a variety of forms such as zoning in the US and green belt in the UK. Since land use regulation has a wide range of economic impacts on housing markets, labour supply, and local environment (Mayer and Somerville 2000, Glaeser and Kahn 2004, Mills 2005, Saks 2008, Gyourko and Molloy 2015, Hilber and Vermeulen 2016), many researchers have explored the origin and determinants of it, but mainly in the context of developed countries. Previous studies find that land use regulation is determined by the incentives and actions of agents in local communities. Homeowners, politicians, and developers all participate in the designation process, and restrictive land use regulation is implemented to protect home value, reduce local dis-amenities, prevent low-income households from moving in, and maintain local fiscal advantage (Fischel 1987, Bates and Santerre 1994, Pogodzinski and Sass 1994, Glaeser and Ward 2009, Hilber and Robert-Nicoud 2013, Been *et al.* 2014). However, the literature to date contains few attempts on investigating land use regulation in the context of developing countries due to a vague understanding of local politics and a lack of comprehensive datasets.

This paper sets out to understand the determinants of land use regulation in China, a developing country that experiences a rapid process of urbanization during the past decades.[35] I investigate the designation process of floor area ratio (FAR) limit, a major form of land use regulation that specifies construction density in China. FAR limit regulates the maximum ratio of the floor area within the proposed property relative to the size of the land parcel. It has a crucial impact on land value and housing supply, as it determines the number of dwellings to be built out by developers. FAR limit also

[35] The share of the urbanized population in China rises from 25.8% in 1990 to 57.4% in 2017 (National Bureau of Statistics in China).

affects local environment because high construction and population densities are associated with negative externalities such as less sunshine, more congestion, and more pollution (Duranton and Turner 2018, Borck and Schrauth 2019, Carozzi and Roth 2020). Therefore, understanding the determinants of FAR limits can provide additional insights into house price dynamics and the urban environment.

Exploring the process of FAR design also has important policy implications. Between 2005 and 2017, the mean value of the FAR limit for residential use is much lower in superstar cities such as Shanghai (1.91) and Beijing (2.27) compared with the national mean level (2.78) in China.[36] This is contradicted with the common sense that superstar cities construct high-rise buildings and benefit from the agglomeration economies. Conversely, cities in the less developed middle and western regions set relatively high FAR limits for residential use. These cities are not experiencing economic prosperity as Beijing and Shanghai. High-rise buildings are constructed and then left vacant, which is widely covered by the media as the 'ghost town' phenomenon.

To understand the determinants of FAR limits in China from a political economy perspective, I develop a spatial equilibrium framework with the promotion incentive of local officials. The model implies that FAR limit is the outcome of local governments trading-off between the benefits (more housing supply, land revenue, and public good provision) and the costs (negative externalities) of high construction density. Local governments with sufficient budgetary revenue are less relied on land sales to provide public goods and opt to design relatively low FAR limits to reduce the negative externalities caused by density. Conversely, cities with fewer budgetary revenue are more financially relied on land sales and design higher FAR limits to generate more fiscal income. The model also shows that the designation process of FAR limit reduces local housing supply and contributes to the affordability issues in some cities.

To test for the theoretical framework's proposition, I exploit a comprehensive dataset of over 200,000 residential land transactions in China between 2005 and 2017 and a county-level panel to study the impact of local budgetary revenue on FAR design. Budgetary revenue is a commonly used measure of local fiscal capacity in China and

[36] Estimates using this paper's baseline sample. See section 3.1 for details.

includes taxes, administration fees, and the shared profits from state-owned enterprises. To mitigate the endogeneity concerns of reverse causality and local confounding characteristics, I first create spatial grids across the country and compare land parcels within a small geographic unit. I then exploit the exogenous variation generated by a central government administrative adjustment policy for identification. The policy turns self-governed counties into prefecture-governed districts, which breaks administrative boundaries, leads to infrastructure improvement, and boosts local agglomeration economy and budgetary revenue. I then utilize a propensity score matching (PSM) approach to mitigate potential selection bias of the treated cities. In line with the theoretical framework, this paper's more credible PSM-IV estimate suggests that a one standard deviation increase in local budgetary revenue will decrease FAR limit by 0.6, which is 44% of the standard deviation of the FAR limit in the baseline sample. I also conduct a spatial boundary design and a placebo test with 1,000 randomly generated treatment dates to address the concerns of unobserved local characteristics and the spurious documentation of the treatment effect.

This paper relates to the literature that explores the economics of land use regulations, including the welfare analysis of land use regulations (Cheshire and Sheppard 2002, Turner *et al.* 2014), the determinants of housing supply restrictions (Glaeser and Ward 2009, Saiz 2010, Hilber and Robert-Nicoud 2013, Been *et al.* 2014) and the consequences (Mayer and Somerville 2000, Gyourko and Molloy 2015, Hilber and Vermeulen 2016). Previous studies have mainly discussed land use regulations in the context of developed countries. A predominant theory proposes that homeowners (or the 'not in my backyard residents', NIMBYs) oppose local new developments and vote for politicians who can introduce restrictive land use regulations to protect their home value. The literature also discusses the fiscal incentive (Rolleston 1987, Bates and Santerre 1994) and the exclusion incentive (Pogodzinski and Sass 1994) of restrictive land use regulations. This paper contributes to the literature by documenting the determinants of land use regulation in the context of a developing country. Theoretically, this paper links the Rosen-Roback spatial equilibrium framework with the politician tournament theory in China (Li and Zhou 2005) and discusses how local officials' incentive and the Chinese 'Land Finance Model' influence land use design. Empirically, previous studies mainly measure land use regulations aggregated at some geographical levels such as The Wharton Residential Land Use Regulatory Index

(Gyourko *et al.* 2008). This paper uses a unique and comprehensive dataset of individual land transactions in China to measure time-varying regulatory restrictiveness at a detailed land plot level. This micro-level dataset also allows me to conduct a rigorous identification strategy.

This paper also relates to the literature on the economics of construction density control in the contexts of both developed countries (Barr 2013, Ahlfeldt and McMillen 2018) and developing countries (Fu and Somerville 2001). Cai, Wang, and Zhang (2017) estimate a dataset of land parcels matched with residential projects and find that developers tend to violate FAR restrictions in more desirable locations in China. Brueckner *et al.* (2017) show that the elasticity of land price with respect to the FAR limit could be a measure of local regulation stringency. Using a national sample, they estimate the elasticity to be city-specific, which shows variation in the stringency of FAR regulation across Chinese cities. However, the literature to date contains few attempts on understanding the determinants of these density control regulations, and this paper aims to fill the gap.

This paper also refers to the discussion of urban density, agglomeration, and negative externalities. While density leads to higher productivity (Duranton and Puga 2014), it also causes air pollution (Carozzi and Roth 2020) and potential damage to the ecosystem (Glaeser and Kahn 2004). This paper contributes to the literature by providing an original political economy story on how local governments trade-off between the benefits and the costs of construction density control to achieve a desirable outcome.

In the end, this paper relates to the literature on Chinese economy including fiscal policies and fiscal decentralization (Jin *et al.* 2005, Han and Kung 2015), urban expansion and career incentive of city leaders (Wang *et al.* 2019), and risks of housing markets in China (Wu *et al.* 2012, Wu *et al.* 2016). This paper contributes to the literature by studying how local governments design land use regulations under the Chinese fiscal system and thus enriches the discussion on local public finance and housing markets in China.

The rest of this paper is structured as follows. Section 2 describes the local fiscal system and land use regulation design in China and provides a theoretical framework to guide the empirical analysis. Section 3 contains the data sources, descriptive statistics, identification strategy, and empirical results. Section 4 concludes.

2. Background and Theoretical Framework

2.1. Institutional Background

2.1.1. Local Fiscal System and Land Auction Market in China

During the past decades, China has experienced several waves of reforms with the aim of fiscal decentralization. Fiscal decentralization was first accomplished in the early 1980s through a fiscal contract system. Under this system, local governments could keep almost all extra revenues generated beyond their pre-set contract responsibilities. Following a major tax reform in 1994, which weakened the budgetary revenue for local governments, and a major housing reform in 1998, city leaders learned that selling land leases was an effective way to generate fiscal revenue. Land revenue has since then become a key feature of local public finance in China (Cao *et al.* 2008). It is classified as 'extrabudgetary revenue' and local governments are not required to share it with the central government. Over the past decades, local governments have increasingly relied on selling land parcels as a major source of fiscal revenue to finance local public goods and infrastructure investments. As shown in Figure 1, land sale revenue equals between 36% and 65% of local government's budgetary revenue in China from 2003 to 2017. As discussed above, budgetary revenue is a commonly used measure of local fiscal capacity in China and includes local taxes, administration fees, and the shared profits from state-owned enterprises, but it doesn't include land sales. Figure 1 also shows that if we take into account the 'indirect' land revenue such as the land appreciation tax, the aggregated land revenue will even exceed local government's budgetary revenue in 2010.

By law, all urban lands are owned by the state in China. Since 1988, the prefecture land bureau has gotten the authority to allocate the use rights of vacant urban lands. The maximum terms of the land use rights are 70 years for residential use, 50 years for industrial use and mixed use, and 40 years for commercial use. In 1990s, most land leases were allocated through negotiation between local governments and developers. In order to control for the corruption occurred during negotiation, the Ministry of National Land and Resources banned negotiated land deals after August 2004. Since then, all urban land leases for private development have been allocated through public auctions. Land auctions are held by local government's land bureau, and detailed information of land parcels are required to be available to the public. According to Cai, Henderson, and Zhang (2013), more than 95% of land auctions were conducted via either English auctions or two-stage auctions. Local governments collect land revenue from these auctions and the land sale serves as an important source of local fiscal revenue besides taxes and administrative fees.

2.1.2. FAR Design and The Objective of Local Officials

Local government's urban planning bureau designs land use regulations such as FAR limit and the share of green area for each land plot to be released. These plots will then be turned over to the land bureau for auction. In practice, based on local governments' documentations and my interviews with local officials and developers, the designation process of FAR limit is mainly through discussion and negotiation between county-level governments and prefecture-level governments. County-level governments propose land use plans to the prefecture-level governments, and the decisions will be made by prefecture-level governments based on different environmental, economic, and urban planning criteria. This paper uses the county-level budgetary revenue in the main empirical analysis because county-level governments, especially the more rural ones ('xian'), usually have a major influence on land use regulation design. This paper also applies the prefecture-level budgetary revenue as a robustness check and the results are reported in the appendix.

Local governments design both an upper bound and a lower bound for FAR limit. This paper defines FAR restriction as the upper bound constraint because the upper limit is always binding in practice and lower bound cases are very rare (Cai *et al.* 2017).

Local governments consider different factors when designing FAR limit: First, there is an 'environmental impact' of construction density. High-rise buildings usually accommodate dense population and are associated with negative externalities such as bad view, less sunshine, more congestion and more pollution. Figure 2 presents an example of the negative externalities caused by density. I compute the average FAR limit within 5km away from 482 air quality stations in China to study the correlation between air pollution and construction density at the local level. The figure shows a positive correlation between neighbourhood construction density and air pollution measured by NO2 emission. Second, high FAR limit can significantly increase land value, as developers are allowed to build out many dwellings upon the land plot. Figure 3 illustrates the positive correlation between FAR upper limit and land price per square meter using over 200,000 residential land transactions in China. Land sale is a crucial source of fiscal revenue for many local governments in China, and higher FAR design can generate more fiscal revenue for local public good provision and infrastructure investment. Third, high FAR limits can increase the supply of housing units, bring down housing costs, and make housing more affordable.

Local governments trade-off between the benefits and the costs of high FAR design to achieve their objectives. City leaders in China have an incentive to pursue for economic growth during their term time. Li and Zhou (2005) find that local officials are more likely to be promoted if provinces experience economic prosperity under their governance. As land sale accounts for a significant proportion of local fiscal revenue, if local leaders only care about raising fiscal revenue to invest in infrastructure projects and boost local GDP, they will design FAR limit as high as the developer’s optimal construction density to maximize land value. Figures 4 and 5 report the budgetary revenue and the average FAR limit (weighted by land plot size) for residential use at the prefecture level in China, respectively. Figure 4 illustrates that cities along the southeast coast have more budgetary revenue compared with the inland cities, as these cities are more economically developed and attract more high-technology companies and high-skilled labours. Figure 5 then shows that cities along the southeast coast tend to design lower FAR limits for residential use, and cities in the less developed central and western regions set higher FAR limits. Figures 4 and 5 together suggest that at least in the more economically advanced coastal cities, local officials consider factors more than just maximizing land sale revenue when designing FAR limits. In fact, if local governments can collect sufficient fiscal revenue from other sources, they tend to design relatively low-rise buildings. Figure 6 shows a clear negative correlation between FAR limits and local budgetary revenue using over 200,000 residential land transactions. To understand the designation process of FAR limits and the trade-off faced by local governments, I propose a spatial equilibrium theoretical framework and the details are discussed in section 2.2.

2.1.3. Turning Counties into Districts

To mitigate the endogeneity concerns of this paper's baseline estimation, I exploit the exogenous variation of local budgetary revenue generated by a central government administrative adjustment policy named 'Turning Counties into Districts' (TCID). The details of this policy are discussed as below.

China has established a unitary centralized power system since 1949. The system of Chinese local administrative division has four levels (from top to bottom): provincial-level, prefecture-level (city-level), county-level, and town-level. As shown in Figure 7, the county-level administration consists of municipal districts, which are more urban and directly governed by prefecture-level governments, and counties, which are more rural and have a higher degree of administrative autonomy in different aspects such as fiscal budget and land supply. This administrative autonomy introduces more flexibility for county leaders to adjust policies based on local economic conditions. However, it also causes administrative boundaries and inefficiencies among different levels of governments. For instance, if a prefecture government wants to implement a city-wide subway network, the county government might oppose and delay this project because the subway station will generate noise and pollution to the county residents.

During the past decades, there is a rapid process of urbanization in China, and many rural counties have been turned into municipal districts to be directly governed by the prefecture-level governments. The major aim of the TCID policy is to boost local economic development by breaking administrative boundaries and promoting cooperation among different levels of governments. In most cases, the TCID policy is conducted following the steps as below: prefecture-level governments first investigate the counties to be adjusted and cooperate with county-level governments to prepare for an administrative adjustment proposal. They then submit the proposal to the provincial government and the state council (the central government). The central government reviews the adjustment plan and make their policy decision based on a variety of local economic and social conditions. In 1993, the ministry of civil affairs from the state council published the criteria that both the prefecture and the county to be adjusted need to satisfy to get the approval of TCID. The criteria include the lower limit of population, the upper limit of employment in the agricultural sector, and some requirements on urban expansion, local budgetary revenue, and industry composition.

From 2000 to 2019, 105 prefecture-level cities in China have turned their counties into municipal districts. As Figure 8 shows, there are two major waves of the administrative adjustments, starting in the early 2000s and the early 2010s respectively. While the first wave is largely driven by central government's instruction, the second wave mainly reflects the demand from the local government side. During the second wave, prefectures actively apply for turning their counties into municipal districts to avoid geographical and administrative obstacles for future urban development. Figure 9 shows a substantial spatial variation in cities that implemented the TCID adjustments between 2000 and 2019. Figure 10 then presents an example of the TCID policy. In June 2002, a prefecture-level city Xi'an got the approval that its Chang'an county (one of the light blue polygons) could be turned into a municipal district (dark blue polygon). After this adjustment, Chang'an district would be directly governed by the prefecture-level government.

The impacts of the TCID policy have been widely discussed. TCID policy usually benefits the prefecture-level government by bringing in extra fiscal revenue from the county-level administration and allowing the prefecture to implement city-wide infrastructure projects. For instance, Foshan turned 4 of its counties into municipal districts in 2002. After the administrative adjustment, Foshan government spent 10 billion RMB on an infrastructure project to connect the 4 newly adjusted municipal districts with the pre-existing central districts. This project significantly reduced transportation costs and led to an industry upgrading in the pre-existing municipal districts because high-skilled workers and high-end industries would concentrate in the central area after all districts were well connected.

However, the TCID policy seems to be a double-edged sword for the rural county to be adjusted. On one hand, the county can benefit from having access to the prefecture-level public goods after the adjustment. On the other hand, the county needs to transfer a large proportion of its fiscal revenue to the prefecture-level government and might compromise to the prefecture-level infrastructure plan. Some county residents are worried that after the TCID policy, more resources will be reallocated from the newly adjusted districts to the pre-existing central districts, because the prefecture officials might have a preference on the central area. Figure 9 presents the night lights in Xi'an before and after the TCID policy, respectively. The figure shows that although the newly adjusted Chang'an district were more urbanized after the TCID policy, the nightlight in the pre-existing central municipal districts became much brighter after the adjustment. This paper assumes that the TCID policy will generate an exogenous increase in the pre-existing central district's budgetary revenue due to the infrastructure improvement, industry upgrading, and the growing agglomeration economies after the adjustment. The details of using the TCID policy as an instrument for local budgetary revenue are discussed in section 3.2.2.

2.2. Theoretical Framework

In this sub-section, I develop a theoretical framework, a static spatial equilibrium model of local governments, households, firms, and developers, to guide the empirical analysis. The model illustrates how local governments trade-off between the benefits and the costs of high FAR design. The model is built on the spatial equilibrium framework developed by Rosen (1979), Roback (1982), and Diamond (2017), and I extend the classic framework by introducing the political incentive of local officials in China (Li and Zhou 2005). The set of players and the timing of the game are as follows: local governments simultaneously choose a FAR limit, sell land parcels, and spend all fiscal revenue including budgetary revenue and land revenue on the provision of local public goods. Households then make location decision among cities based on their expected utility level and the payoffs are realized. In the end, the urban system reaches a spatial equilibrium, and households settle in one city and have no incentive to move.

2.2.1. Firms

Suppose that homogenous firms use 1 unit of capital and L_F units of labour to produce tradable goods. The production of tradable goods Y follows a simple Cobb-Douglas function:

$$Y = A \times L_F^{\alpha}$$

Where A stands for the total factor productivity (TFP) and L_F stands for labour input. Suppose that $0 < \alpha < 1$, the unit cost of labour is w, and the unit cost of capital is k. The unit cost of capital is determined by the national credit supply and demand conditions and is thus exogenous. To maximize profit, each firm will use labour L_F:

$$L_F = (\frac{w}{\alpha A})^{\frac{1}{\alpha - 1}}$$

Suppose the price of tradable goods is constant at 1. Under the assumption of perfect competition and free entry and exit of firms, profit π_F equals to 0. Therefore, wage w is determined as follow:

$$\pi_F = A(\frac{w}{\alpha A})^{\frac{\alpha}{\alpha-1}} - w\left(\frac{w}{\alpha A}\right)^{\frac{1}{\alpha-1}} - k = 0$$

$$w = \alpha^{\frac{1}{\alpha}} k^{\frac{\alpha-1}{\alpha}} \mathrm{A}^{\frac{2-\alpha}{\alpha}} \tag{1}$$

Equation (1) suggests that wage w is exogenously determined by local TFP A, α, and capital cost k.

2.2.2. Households

Suppose that there is an urban system which consists of multiple cities. Homogenous households can move across cities with no migration cost and make their location decision based on the expected utility level. Household's utility is determined by the consumption of housing q_h, the consumption of tradable goods q_c, public goods g provided by local government, and the negative externalities e that are associated with construction density such as congestion and pollution. Public goods are equally shared by all the population L living in the city. Suppose that p_h represents housing price per square meter. The utility function and budget constraint for households are thus as follow:

$$U = q_h^{\beta} q_c^{1-\beta} + \left(\frac{g}{L}\right)^{\gamma} - e \tag{2}$$

$$s.t. \quad w = p_h q_h + q_c$$

Where $0 < \beta < 1$ and $0 < \gamma < 1$. To maximize individual utility level, each household will consume housing q_h and tradable goods q_c as below:

$$\frac{q_h}{q_c} = \frac{\beta}{1-\beta}$$

Let d denote housing stock within a city and suppose that local government releases N land parcels with size S and FAR upper limit f to the housing market. Under the assumption of housing market clearing:

$$q_h L = NSf + d$$

Housing consumption q_h and tradable good consumption q_c are thus determined as below:

$$q_h = \frac{NSf+d}{L},\ q_c = \frac{1-\beta}{\beta}\left(\frac{NSf+d}{L}\right)$$

2.2.3. Land Markets and Developers

Suppose that there are identical developers within a city. Developers purchase land parcels from the local government. Let r denote the land price per square meter and $c(f)$ denote the construction cost per square meter. $c(f)$ is a convex function with respect to f ($c'(f) > 0, c''(f) > 0$) because the marginal construction cost will increase as the building height increases. Developers bid for land parcels based on their expected house price p_e and the construction cost $c(f)$. After acquiring land plots, developers will build projects with construction density f.[37] Developer's profit π_d is thus given by:

$$\pi_d = Sfp_e - Sr - Sc(f)$$

Under the assumption of perfect competition and free entry and exit, developers make zero profit and land price r is given by:

$$r = fp_e - c(f)$$

Suppose that housing market is clear before local government releasing any new land parcels. Let d denote housing stock and L_0 denote local population before land release. Total housing demand equals total housing supply:

$$q_h L_0 = d$$

Developers expect house price p_e to be at the same level as what they observe before any new land plot is released. The expected housing price p_e and land price r are determined as below:

$$p_e = \frac{\beta L_0 w}{d}$$

$$r = \frac{\beta f L_0 w}{d} - c(f)$$

This paper assumes that r is an increasing and concave function with respect to f ($\frac{\partial r}{\partial f} > 0, \frac{\partial^2 r}{\partial^2 f} < 0$), meaning that FAR limit has a positive impact on land price per

[37] This assumption is in line with Cai, Wang, and Zhang (2017)'s finding that the upper FAR limits are always binding for residential projects in China.

square meter, but the marginal effect is decreasing as FAR limit increases.[38] This assumption is plausible because: Frist, developers can build and sell more housing units as FAR limit increases, so FAR limit is likely to have a positive impact on land value. Second, land value will not increase infinitely because construction cost $c(f)$ will also increase with the building height, suggesting that the second-order derivative $\frac{\partial^2 r}{\partial^2 f}$ is negative. This assumption is also in line with previous findings on the positive correlation between FAR limits and land value (Brueckner *et al.* 2017) and on the binding FAR upper limits (Cai *et al.* 2017).[39]

2.2.4. Negative Externalities

High population density is associated with negative externalities such as congestion and pollution (Duranton and Turner 2018, Borck and Schrauth 2019, Carozzi and Roth 2020). High construction density f also leads to bad views and less sunshine. These negative externalities will adversely affect household's utility.

Let $\frac{NSf+d}{NS+S_0}$ denote the overall construction density within a city, where S_0 denotes the land area of housing stock d, and NS denotes the area of new land supply. $e(\frac{NSf+d}{NS+S_0})$ denotes all the negative externalities caused by density and is defined as a convex function with respect to f (therefore, $\frac{\partial e(\frac{NSf+d}{NS+S_0})}{\partial f} > 0, \frac{\partial^2 e(\frac{NSf+d}{NS+S_0})}{\partial^2 f} > 0$). I simplify $e(\frac{NSf+d}{NS+S_0})$ to $e(f)$ because all the parameters other than f in this function are exogenously determined.

2.2.5. Local Officials Design the Optimal FAR Limits

Suppose that local government simultaneously designs FAR limit f for N land parcels with size S and sell them to developers. Local government then spends all fiscal revenue including land sales NSr and budgetary revenue B on public good provision. Budgetary revenue B includes local taxes and administrative fees and is first treated

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[38] $\frac{\partial r}{\partial f} = \frac{\beta L_0 w}{d} - c'(f)$ and $\frac{\partial^2 r}{\partial^2 f} = -c''(f)$. It is easy to prove that $\frac{\partial^2 r}{\partial^2 f} < 0$. This paper assumes that $\frac{\beta L_0 w}{d} - c'(f) > 0$ so that $\frac{\partial r}{\partial f} > 0$.

[39] Table A3 provides empirical evidence about the positive correlation between land price per square meter and FAR upper limit, especially for residential use, using land transaction data in China.

as exogenous in the model. I will discuss the endogenized budgetary revenue B in section 2.2.7.

This paper assumes that the provision of public good g follows a simple production function with government's labour input normalized to 1 and productivity A_G:

$$g = A_G(NSr + B) \tag{3}$$

Let L denote population within a city. Under the assumption of local housing market clearing, housing supply equals housing demand:

$$NSf + d = q_h L \tag{4}$$

Households make location choices based on their expected utility level as suggested in equation (2). I plug equations (3) and (4) into equation (2) to rewrite the utility function for each household:

$$U = \overbrace{\frac{\left(\frac{1-\beta}{\beta}\right)^{1-\beta}(NSf+d)}{L}}^{supply\ effect} + \overbrace{\left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma}}^{fiscal\ effect} - \overbrace{e(f)}^{externality\ effect} \tag{5}$$

Equation (5) suggests that FAR limits influence a household's utility level through three channels. First, when local government sets higher FAR limits, there will be more supply of housing units, which bring down housing price and increase the quantities of housing consumed by households (supply effect). Second, higher FAR limits generate more land sale revenue and enable local governments to provide more public goods (fiscal effect). Third, there is a negative externality effect associated with construction density, which is represented by $e(f)$. Figure 11 illustrates the intuition of equation (5): The red and blue curves represent the two positive effects of high FAR design, and the black curve represents the negative externalities associated with density. All effects will increase as the FAR limit increases, but at different paces. If local government behaves as a benevolent social planner and only cares about local resident's utility level, they will set the optimal FAR limit at a point where the positive curves and the negative curve have the largest gap.

This paper assumes an 'open city' scenario, meaning that households can move freely across cities to achieve the highest utility level and there is no migration cost. When the urban system reaches a spatial equilibrium, every household will have the same utility $\overline{U}$ and no incentive to move out. The population L within a city is thus endogenously determined as follow:

$$\overline{U} = \frac{\left(\frac{1-\beta}{\beta}\right)^{1-\beta}(NSf+d)}{L} + \left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma} - e(f) \tag{6}$$

Equation (6) suggests that the population size within a city will increase if there are more housing supply and public goods and will decrease if there are more negative externalities caused by density.

Suppose that the local labour market is cleared. The aggregate economic output within a city, Y, is represented by the sum of all firms' outputs:

$$Y = L \times (\frac{w}{\alpha A})^{\alpha-1} \times A \times (\frac{w}{\alpha A})^{\frac{\alpha}{\alpha-1}}$$

Based on the 'politician tournament theory' (Li and Zhou, 2005), city leaders in China are motivated to boost local economy so that their probabilities of being promoted will increase. This paper thus assumes the optimal FAR limit f^* for local officials will maximize the aggregate economic output Y as well as the population size L[40] within a city:

$$f^* = \underset{f}{arg\,max}\,(L)$$

This paper then proves the following inequality:[41]

$$\frac{\partial f^*}{\partial B} < 0 \tag{7}$$

This inequality illustrates the main proposition to be tested in this paper:

Proposition 1 – Local governments with more budgetary revenue opt to design lower FAR limits in order to reduce the negative externalities caused by density and to maximize local GDP and population size.

2.2.6. FAR Limit, House Price, and Housing Consumption

This paper then explores the impact of FAR design on housing markets. Under the assumption of housing market clearing, housing price is determined as below:

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[40] The objective of local officials to increase population size is supported by the competition among local governments to attract young talents in China ('qiang ren da zhan'). Local governments provide a series of benefits to undergraduates and postgraduates who decide to settle in. These benefits include a relaxation of the hukou requirements and local housing subsidies.
[41] See proof in Appendix C.

$$p_h = \frac{wL}{NSf + d} - \frac{1-\beta}{\beta}$$

This equation suggests that demand-side factors such as population and wage will increase house prices, and supply-side factors (dwelling stock, FAR limit, number of land plots, land area) will reduce house prices. I then prove the following inequalities:[42]

$$\frac{\partial L}{\partial B} > 0,\ \frac{\partial p_h}{\partial B} > 0 \qquad (8)$$

These equalities predict that: First, cities with more budgetary revenue will attract more population. Second, these richer cities also have higher housing prices, which are driven by both demand-side factors (more public goods and population) and supply-side factor (fewer housing units caused by lower FAR design).

Under the assumption of housing market clearing, housing consumption is determined as below:

$$q_h = \frac{NSf + d}{L}$$

It is easy to prove that:[43]

$$\frac{\partial q_h}{\partial B} < 0$$

Inequality (8) illustrates the second hypothesis to be tested in this paper:

Proposition 2 – Cities with more budgetary revenue have higher housing prices. This is driven by both demand-side factors (more population and public good provision) and supply-side factors (lower FAR limits and fewer housing units).

2.2.7. Endogenous Local Budgetary Revenue

So far, this paper assumes an exogenously determined budgetary revenue B. In this sub-section, I endogenize local budgetary revenue by simply defining it as a local income tax:

$$B = L\tau w$$

Where τ denotes tax rate. This paper then proves that:[44]

[42] See proof in Appendix C.
[43] See proof in Appendix C.
[44] See proof in Appendix C.

$$\frac{\partial f^*}{\partial A} < 0$$

This inequity predicts that cities with higher TFP will collect more budgetary revenue and design lower FAR limits by following the same mechanism as discussed above.

3. Empirical Analysis

The main purpose of the empirical analysis is to test for propositions (1) and (2) and study the impact of local budgetary revenue on FAR design and house price. The details of the empirical analysis are discussed as below.

3.1. Data and Descriptive Statistics

This paper's main estimation sample uses 202,816 residential land transactions in 281 prefecture-level cities and 1,804 counties in China from 2005 to 2017. The data source is the official website of China land market, which covers the vast majority of land transactions in China. The main dataset records detailed information at the plot level including land transaction price, address, the date of transaction, the upper and lower limits of FAR, the type of land use, a land quality evaluated by government, land area, planned total floor area, the type of auction, land use, and the land bidder. This paper defines FAR restriction as the upper bound of the FAR limit, because the cases of lower bound constraints are very rare, and the upper limit is almost always binding (Cai, Wang, and Zhang 2017). I use Gaode Map API to geo-code all the land parcels based on their location information. As Figure 12 shows, the geocoded land parcels cover most major cities in the country. Besides, the land parcels are widely spread within cities. For instance, Figure 13 presents a rich amount of geocoded land transactions both in the central area and at the urban fringe of Beijing.

This paper first identifies the land use of each plot based on its planning description and then selects residential land transactions for the main empirical analysis. Residential land sale serves as the major source of land revenue and accounts for over 75% of all the land sales in this paper's estimation sample. Wang, Zhang, and Zhou (2019) also document that about three quarters of the land sale revenue created through public auctions come from the sale of residential land. This paper estimates a sample including commercial and industrial lands and the results are reported in the appendix.

As discussed before, I utilize a spatial grid approach to control for time-invariant local characteristics. I create a fishnet that covers all the land transactions in the baseline sample and the size of each grid is 3km × 3km. Figure 13 shows the spatial grids in Beijing. By controlling for grid fixed effects in the main specification, this paper compares land parcels within relatively small geographical areas and mitigates the concern of unobserved time-invariant local features such as historical construction density, geographical obstacles, and local amenities.

This paper collects nation-level, prefecture-level, and county-level characteristics from different sources including China Financial Statistical Yearbook, China Financial Statistics of Cities and Counties, City Statistical Yearbook, local statistical yearbooks, and local government statistical reports. The administrative adjustment records are collected from the Ministry of Civil Affairs. I merge the land parcel data with the county-level and prefecture-level panels to construct the baseline estimation sample. Following the literature, I dropped the top 1% and the bottom 1% of observations in terms of FAR restriction to mitigate the bias caused by extreme values. The key explanatory variable, budgetary revenue of local government, is standardised so that I can easily interpret the estimated coefficients. This paper subtracts the sample mean of budgetary revenue from itself and divide this difference by the standard deviation. This transformation allows me to interpret the estimated coefficient as an increase in the FAR limit due to a one standard deviation increase or decrease in local budgetary revenue. This paper also collects station-level air quality data from China National Environmental Monitoring Centre.

Basic summary statistics computed for a sample of residential land transactions in China from 2005 to 2017 are detailed in Panel A of Table 1. There are in total 202,816 residential land transactions. The average value of the land transaction price is 53 million RMB (around 6 million GBP[45]), and the average size of the land parcel is 25,000 square meters. The key land use regulation explored in this paper, FAR upper limit, has a mean value of 2.8 and a standard deviation of 1.4. As Figure 14 shows, most land parcels have FAR restrictions between 1 and 6, and there is significant bunching at round numbers. The mean value of distance to CBD is 42 km. Panels B and C of Table 1 then shows the descriptive statistics for city-level and county-level

[45] Based on the currency exchange rate in November 2020.

characteristics from 2005 to 2017, respectively. The key explanatory variable in the empirical analysis, budgetary revenue at the county level, has a mean value of 1,313 million RMB.

Figures 4 and 5 present the spatial patterns of local budgetary revenue and FAR design at the prefecture level in China, respectively. Figure 4 suggests that regional core cities and cities along the southeast coast tend to have more budgetary revenue. These cities, including Tier 1 cities such as Beijing and Shanghai, are reckoned as the more economically advanced cities. Conversely, Figure 5 presents the average FAR limit for residential use and shows a reverse spatial pattern. I compute the weighted average FAR limit[46] at the prefecture level using residential land plots between 2005 and 2017, and the map suggests that regional core cities and cities along the southeast coast tend to set relatively low FAR limits for residential use. Figures 4 and 5 together suggest that cities with more budgetary revenue tend to design lower FAR limits. This stylized fact based on raw data is in line with proposition (1) from the theoretical framework.

This paper also estimates a sub-sample of land transactions with information about the monopsony power of land bidders to study the impact of developer's bargaining power on FAR design. The results are discussed in section 3.4.

3.2. Empirical Specifications and Identification Strategy

3.2.1. Main Specification and Endogeneity Concerns

This paper's empirical strategy is designed to test for proposition (1) and explore the determinants of FAR limits in China using land transaction data. I first estimate the following equation using OLS:

$$FAR_{icdgym} = \phi_d + \rho_g + \beta Budget_{dy} + \delta_{ym} + \gamma' X_i + \tau' Z_{cy} + \varepsilon_{icdgym} \tag{9}$$

where i indexes individual land parcel, y indexes transaction year, and m indexes transaction month. The key explanatory variable, $Budget_{dy}$, represents the budgetary revenue of county d in year y. A vector of county fixed effects is represented by ϕ_d and a vector of spatial grid fixed effects is represented by ρ_g. δ_{ym} is a set of time dummies (year-month fixed effects) and X_i is a set of land parcel controls including land area, distance to CBD, a land quality measured by government, the type of land

[46] The average FAR limit is weighted by the size of each land plot.

auction, and the longitude and latitude of the land plot.[47] Z_{cy} is a set of prefecture-level time-varying characteristics including population, average salary, local industry composition, the number of universities and the number of hospitals in city c in year y. This paper estimates this equation by OLS, clustering standard errors at the grid level to account for potential spatial autocorrelation in FAR design and local housing market conditions. The parameter of interest is β, measuring the impact of local budgetary revenue on FAR restriction. This paper also estimates a similar specification as equation (9) but replaces county-level budgetary revenue with prefecture-level budgetary revenue to test if the impact is robust at a higher administration level.

One important caveat with the OLS estimates of equation (9) is that the explanatory variable $Budget_{dy}$ is likely endogenously determined, causing the estimate to be biased. There are two major concerns. First, since FAR limit is correlated with land value, and certain types of local taxes such as land appreciation tax and stamp duty are computed based on land price, FAR limit is likely to have a direct effect on local government's budgetary revenue, which leads to reverse causality. This simultaneity issue will underestimate the negative impact of local budgetary revenue on FAR limit.[48] Second, unobserved local features might cause bias in the OLS estimate. For instance, the population density in the pre-existing informal housing upon the land plot will increase the resettlement costs for land acquisition (Fu and Somerville, 2001), and local governments might design high FAR limits to compensate for the increasing acquisition costs. Meanwhile, the literature suggests that there is a positive impact of informal housing on accommodating migrant inflows within cities (Niu *et al.* 2020), so the density of informal housing might have a positive effect on local budgetary revenue. This confounding factor is not fully controlled for in the main specification due to data availability and will underestimate the negative effect of budgetary revenue on FAR limits.[49] Besides, the literature suggests that there is a significant amount of corruption in the land auction market in China (Cai, Henderson, and Zhang 2013), and the time-varying local corruption level might be correlated with both FAR design and budgetary revenue. Lastly, previous research suggests that land use regulations tend to

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[47] I control for the longitude and latitude of each land plot to take into account the requirement of sunshine time for buildings, as locations in different longitudes and latitudes have different sunshine angles and exposure.
[48] See proof in Appendix D.
[49] See proof in Appendix D.

be historically dependent: if there are many low-rise buildings within a neighbourhood, local officials are more likely to design low FAR limit for a newly released land parcel there.

3.2.2. Identification Strategy

To address the endogeneity concerns as discussed above, this paper first creates 3km × 3km spatial grids covering the whole country and applies a grid fixed effect strategy to compare land parcels within a small geographic unit. This method allows me to control for time-invariant local features such as historical construction density and geographical obstacles. Because of the rich land transactions in the sample, I can observe sufficient variations in FAR limits after controlling for the grid fixed effects. I also control for the effect of county-level time-invariant features and macro trends by including county fixed effects and year-month fixed effects, respectively. To mitigate the concern of high-skilled labour sorting into superstar cities, I control for time-varying local income level and amenities such as average salary, industry composition, the numbers of universities and medical facilities.

However, two potential endogeneity issues remain after I control for multiple fixed effects and the time-varying local characteristics: First, FAR limit is likely to be correlated with local budgetary revenue through taxes related to land value. Second, unobserved time-varying local factors such as corruption and the density of informal housing are likely to affect both FAR design and budgetary revenue.

To address these endogeneity concerns, this paper proposes an instrument variable strategy by exploiting the exogenous variation generated by TCID, a central government administrative adjustment policy.[50] Figure 8 shows that the administrative adjustments are widely implemented across Chinese cities and provides sufficient spatial variations for the empirical analysis. The identification assumption is that pre-existing central municipal districts are likely to be the 'winner' of this policy and can collect more budgetary revenue after the adjustment, because TCID policy breaks administrative boundaries and stimulates infrastructure improvement in the centre area. Meanwhile, there is no direct correlation between local FAR design and the implementation of TCID policy, as the adjustment decision is based on certain criteria set by the central government and not likely to be influenced by local land use

[50] The details of this policy are discussed in section 2.1.3.

regulations. The administrative adjustment might directly influence land plots within the newly adjusted districts because these districts will be governed by the prefecture-level government with a different planning idea. This paper thus removes all the 'new districts' in the estimation sample and the treatment group will only include the land plots within the pre-existing central municipal districts. This paper then estimates the following first stage regression:

$$Budget_{dy} = \beta TCID_d \times Post_y + Controls + \varepsilon_{dy} \tag{10}$$

where d indexes individual county/district and y indexes year. The instrument, $TCID_d \times Post_y$ is an indicator variable that takes the value of 1 if d is a pre-existing municipal district within a prefecture that gets the TCID approval and y is after the implementation of the administrative adjustment in the prefecture. I then follow a two-stage-least-square (2SLS) strategy to estimate the impact of the budgetary revenue on FAR design using the budgetary revenue variable instrumented with the TCID policy.

For the instrumental variable estimator to be consistent and unbiased, the conditions are as below: First, the TCID policy affects local budgetary revenue directly (relevance). Second, the treatment is as good as randomly assigned (independence). Third, the policy influences FAR design only through changes in local budgetary revenue (exclusion restriction). This paper proves the instrument relevance by reporting both the first-stage results and the F-statistics. Although Figure 9 presents substantial spatial variations in the TCID adjustments across Chinese cities, it is challenging to ensure both independence and exclusion restriction. This paper argues that the TCID policy does not have a direct correlation with local FAR design because the policy decision is based on certain criteria set by the central government. However, there is an obvious concern about the potential selection bias of the treated cities. Prefectures can get the TCID approvals because these cities are experiencing a rapid process of urbanization and can meet the criteria set by the central government. Some unobserved local trends during the urbanization process might be correlated with both local FAR design and the TCID policy. To address this concern, this paper applies a propensity score matching (PSM) approach. I first estimate a city's propensity to be treated using a logit regression with explanatory variables including the population, population growth rate, industry composition, GDP per capita, and budgetary revenue, which are the criteria that central government uses to evaluate local government's application for the administrative adjustment. Next, I select one counterfactual city in the same year with the propensity score closest to the treated city. These matched cities offer a counterfactual urbanization path for how the treated cities would have experienced, had they not been approved to have the administrative adjustment. From estimating the PSM sample, this paper can mitigate the concern of selection bias and compare cities experiencing similar urbanization process before the TCID policy.

One might argue that even the PSM sample is selected based on the observed local characteristics, and there is still a concern of unobserved local features. I mitigate this concern by using a spatial boundary design and selecting land plots within 5 km away from county boundaries. Land plots close to county boundaries tend to have similar neighbourhood and are highly comparable. This paper then uses these land plots to estimate the impact of local budgetary revenue on FAR design and the results are reported in section 3.5.

3.3. Main Results

Table 2 summarizes the results from estimating equation (9) using a sample of residential land transactions in China between 2005 and 2017. Additional covariates are included into the estimation sequentially.

Columns (1) to (3) report the naïve OLS estimates. Column (1) controls for land parcel characteristics, year-month fixed effects, and county fixed effects. Column (2) includes the spatial grid fixed effects, and column (3) further controls for a vector of time-varying prefecture-level characteristics such as population, industry composition, and average salary. The standard errors in all specifications are clustered at the grid level to allow for a degree of spatial autocorrelation. Columns (1) to (3) show that budgetary revenue has a negative impact on FAR design, and all estimates are statistically significant at 1% level. To quantify the results, the estimate from column (3) suggests that a one standard deviation increase in county-level budgetary revenue will decrease FAR limit by 0.08, which is around 6% of the standard deviation of FAR limits in the baseline sample. The negative coefficient from the OLS specification is in line with proposition (1) that local governments with more budgetary revenue opt to design lower FAR limits.

As discussed in section 3.2.1, potential endogeneity issues might lead to biased OLS estimates. This paper then applies the instrumental variable strategy to study the impact of budgetary revenue on FAR design and the IV results are reported in columns (4) to (6) of Table 2. All the coefficients have the expected signs and are statistically significant at 1% level. The more credible IV estimate in column (6) suggests that a one standard deviation increase in local budgetary revenue will decrease FAR limit by 0.6, which is around 43% of the standard deviation of the FAR limit in the baseline sample. The IV estimates are larger than the OLS estimates in Table 2, which is in line with the expectation that reverse causality and unobserved confounding factors will underestimate the negative impact of budgetary revenue on FAR design.

Regarding the validity of the instrument, the Kleibergen-Paap F statistics in Table 2 suggest that weak instrument is not a concern. In addition, Table 3 reports the first-stage estimation results. The coefficients from columns (1) to (3) are all positive and statistically significant, meaning that as expected, the TCID policy will increase the budgetary revenue of the pre-existing municipal districts.

A PSM-IV method is then applied to mitigate the concern of potential selection bias. Table 4 compares different variables between the treated and the control cities before and after the propensity score matching, respectively. The table shows that there is a significant difference between the treated and the control cities before the propensity score matching: cities that get the TCID approval usually have more population and higher budgetary revenue. These cities are more likely to experience a rapid process of urbanization, which leads to potential selection bias. Table 4 then shows that after the propensity score matching, the treated and the control cities are well balanced regarding different characteristics, either used or not used in the matching, as T statistics are insignificant for all variables. Therefore, cities in the PSM sample are likely to experience a similar urbanization process and more comparable for the empirical analysis.

Figure 15 illustrates the average budgetary revenue for the treated and the control districts before the propensity score matching. Most cities in my estimation sample have implemented the TCID after 2010, and Figure 15 presents a clear gap in local budgetary revenue between the treated and control groups prior to 2010. This leads to the selection concern that cities getting TCID approvals are also urbanizing more rapidly before the introduction of the policy. Figure 16 then presents the time trends of budgetary revenue in the treated and the control districts after the propensity score matching. This figure shows a similar trend in budgetary revenue prior to the implementation of the TCID policy and a significant increase in budgetary revenue of the treated districts after 2010. Figures 15 and 16 together suggest that the propensity score matching can mitigate potential selection bias by selecting a sample of comparable cites with parallel pre-treatment trends. In addition, Figures 17 and 18 show that the treated and the control districts in the PSM sample tend to design similar FAR limits before 2010, and the treated districts start to design lower FAR limits after 2010. This is in line with the expectation that the TCID policy can increase local budgetary revenue in the central districts and thus reduce local FAR limits. I will formally test for the impact of the TCID policy on FAR design in section 3.4.

This paper then re-estimates the impact of budgetary revenue on FAR design using the PSM sample and the corresponding results are reported in Table 5. All the IV estimates in columns (4) to (6) are statistically significant and have the expected signs. The F statistics suggest that weak instrument is not a concern, and the first-stage results reported in Table 6 show that the instrumental variable significantly correlates with budgetary revenue in an expected way. The more rigorous estimate in column (6) of Table 5 suggests that a one standard deviation increase in local budgetary revenue will decrease FAR limit by 0.62, which is around 44% of the standard deviation of FAR limit in the baseline sample.

This paper concludes from these findings that the impact of budgetary revenue on FAR limit is well identified. In line with proposition (1), I find that local budgetary revenue has a negative impact on FAR design. As the theoretical framework shows, local governments trade-off between the benefits (fiscal revenue and housing supply) and the costs (negative externalities) of FAR design. If a local government can collect sufficient budgetary revenue from sources other than land sales, it will put more weight on the negative externalities caused by density and set lower FAR limits. Conversely, local governments with fewer budgetary revenue are more relied on land sales and will design higher FAR limits to raise more fiscal revenue.

3.4. Additional Results

3.4.1. The Impact of TCID Policy on FAR Design

In this sub-section, I directly estimate the impact of the TCID policy on FAR design. Figures 17 and 18 first present both the annual and the quarterly average FAR limits in the treated and the control districts respectively using the PSM sample. Both figures suggest a near-identical FAR trend prior to the administrative adjustment and a significant lower FAR limit in the treated districts after TCID policy. I then explore the quantitative impact of TCID policy on FAR design by estimating a difference-in-difference specification as shown below:

$$FAR_{icdgym} = \phi_d + \rho_g + \beta Post_y \times Treat_d + \delta_{ym} + \gamma' X_i + \tau' Z_{cy} + \varepsilon_{icdgym} \qquad (11)$$

where $Post_y$ is a dummy equalling to zero if year y is prior to the administrative adjustment, and $Treat_d$ is a dummy equalling to one if county d is a pre-existing municipal district and within a prefecture that implements the TCID adjustment. The parameter of interest is β, measuring the impact of the administrative adjustment policy on FAR design.

Table A1 reports the estimation results using both the baseline sample and the PSM sample. In line with the main findings, columns (1) to (6) all report negative and statistically significant coefficients, suggesting that after the implementation of the TCID policy, the pre-existing municipal districts will design lower FAR limits compared with other districts and counties. The more credible estimate in column (6) suggests that the TCID policy decreases local FAR limits by 0.13.

3.4.2. Population Density and Negative Externality

To test for the assumption in the theoretical framework that high density causes negative externalities, I estimate the following equation using city-level panel data:

$$Ln(Y)_{cy} = \phi_c + \beta Ln(Density)_{cy} + \delta_t + \tau' Z_{cy} + \varepsilon_{cy} \qquad (12)$$

where c indexes each city and y indexes time periods. The variable $Ln(Density)_{cy}$ represents the natural logarithm of population density of city c in year y. A vector of city fixed effects is represented by ϕ_c. δ_y is a set of time dummies (year fixed effects) and Z_{cy} is a set of city-level controls such as GDP per capita, average salary, and industry composition. The dependent variable $Ln(Y)_{cy}$ represents two different measures of air quality including PM10 and Air Quality Index (AQI). The parameter of interest is β, measuring the impact of population density on air quality.

The estimated results are reported in Table A2. All the estimated coefficients are positive and statistically significant, meaning that cities with higher population density have worse air quality. This is in line with the assumption in the theoretical framework that high density is associate with negative externalities.

3.4.3. The Impact of Budgetary Revenue on House Price

Proposition (2) from the theoretical framework predicts that cities with higher budgetary revenue will have higher housing prices. This is driven by both demand-side factors (more public goods and population) and supply-side factors (lower FAR limits and less housing supply). To test for proposition (2) and measure the impact of budgetary revenue on housing prices, this paper estimates the following specification using a city-level panel data:

$$Ln(HP)_{cy} = \phi_c + \beta Ln(Budget)_{cy} + \delta_y + \tau' Z_{cy} + \varepsilon_{cy} \quad (13)$$

where c indexes each city and y indexes year. The variable $Ln(Budget)_{cy}$ represents the natural logarithm of budgetary revenue in city c in year y. A vector of city fixed effects is represented by ϕ_c. δ_y is a set of time dummies (year fixed effects) and Z_{cy} is a set of city-level time-varying controls. The dependent variable $Ln(HP)_{cy}$ reflects the natural logarithm of housing price per square meter in city c in year y. The parameter of interest is β, measuring the impact of budgetary revenue on housing prices.

Table A3 summarizes the results by estimating a panel of Chinese cities. All main estimates are positive and statistically significant, suggesting that local budgetary revenue increases housing prices. Compared with the OLS estimates in columns (1) and (2), column (3) reports a larger coefficient after I utilize the TCID as an instrument for local budgetary revenue. Column (4) reports the first-stage result, which is significant and in line with the identification assumption. Consistent with proposition (2), this paper finds a positive impact of budgetary revenue on housing prices.

Figures B1 and B2 then present housing prices and housing affordability (measured by the average house price divided by the average salary) at the prefecture level in China in 2017. Both figures suggest that cities along the south-eastern coast, including superstar cities such as Shanghai, tend to have higher housing prices and face severe housing affordability problems. As Figure 5 shows, these cities also design lower FAR limits due to the trade-off as discussed above. The restrictive land use regulations in these cities reduce housing supply and contribute to the housing affordability issues.

3.4.4. FAR limits for Non-Residential Use

Does local budgetary revenue also influence the FAR design for non-residential use? To answer this question, I first compute the weighted average FAR limits for non-residential use across the country. As Figure B3 shows, there is no clear spatial pattern of the FAR design for industrial and commercial uses. I then re-estimate equation (9) using a sample of non-residential land plots and the results are reported in Table A4. The estimated impact of budgetary revenue on FAR limits for commercial use are all insignificant, and the estimated impact of budgetary revenue on FAR limits for industrial use are statistically significant but marginal.

These estimation results first mitigate the endogeneity concern of unobserved spatial characteristics in the main specification. One might argue that there are fewer land plots available in the more economically developed cities, and the scarcity of land plots might have substitutional impact on FAR limits, as local governments will design high FAR limits given the limitation of horizontal expansion. However, if these cities are indeed concerned about the availability of land plots and opt to design high FAR limits, they should also design high FAR limits for non-residential land parcels. As Table A4 shows, the estimated impact of local budgetary revenue on FAR limits for non-residential use is either insignificant or marginal, suggesting that the main result is not driven by the geographical scarcity in more economically developed cities.

Table A4 also provides supportive evidence for this paper's theoretical framework. The theory predicts a negative effect of budgetary revenue on FAR design for residential use because high FAR limit can increase residential land value and influence the substitution between land sales and local budgetary revenue. If FAR limit doesn't have a strong and positive impact on land value, then the impact of local budgetary revenue on FAR design will also be marginal. Table A5 reports the elasticity of land price per square meter with respect to FAR limit for different land uses. Columns (1) to (3) suggest that while this elasticity is around 43% for residential lands, it becomes much lower for other land uses. Especially for industrial use, the elasticity is still positive but only at around 7%, as the FAR design for industrial use is mainly determined by the technical requirements of manufacturing companies. In this case, the quantity of properties to be built out upon an industrial land plot is not a major consideration for land bidders. Besides, some local governments in China intentionally lower the land price for commercial and industrial use to attract firms and manufacturing companies, which further weakens the fiscal motive of high FAR limit.

3.4.5. Transfer Payment from The Central Government

Another major source of fiscal revenue for local governments in China is the transfer payment from the central government. In this sub-section, I estimate the impact of central government transfer payment on FAR design. The prefecture-level fiscal transfer data comes from China Financial Statistics of Cities and Counties. Since the local public transfer data becomes unavailable after 2009, this paper uses the share of local government's transfer payment in 2007 and the annual national transfer payment trend to estimate yearly transfer payments at the local level. As equation (14) illustrates, $Transfer_{cy}$ represents the estimated transfer payment in prefecture-level city c in year y. $\frac{Transfer_{c2007}}{Total_{2007}}$ represents the share of city c's transfer payment relative to the national level in year 2007, and $Total_y$ denotes the national trend of transfer payment.

$$Transfer_{cy} = \frac{Transfer_{c2007}}{Total_{2007}} \times Total_y \tag{14}$$

This paper then studies the impact of the estimated local transfer payment on FAR limit by estimating a specification similar to equation (9). The results are reported in Table A6. All estimates are negative and statistically significant at 1% level. The most credible estimate in column (3) suggests that a one standard deviation increase in transfer payment will decrease FAR limits by 0.09. This finding is in line with proposition (1) that if local governments can collect sufficient fiscal revenue from sources other than land sales, they will design relatively low FAR limits to reduce the negative externalities caused by density. The estimated effect of transfer payment on FAR design is larger than the OLS estimation of budgetary revenue as reported in Table 2, potentially because transfer payment comes from the central government and is less influenced by local features compared with budgetary revenue. The endogeneity issues as discussed in section 3.2.1 are thus less pronounced in this specification.

3.4.6. Monopsony Power of Developers

In the theoretical framework I assume perfect competition among developers. This sub-section provides additional findings about the impact of land market competition on FAR design. To do so, I estimate the following equation by OLS:

$$FAR_{icdym} = \phi_c + \rho_d + \beta Monopsony_i + \delta_{ym} + \gamma' X_i + \tau' Z_{cy} + \varepsilon_{icdym} \tag{15}$$

where i indexes individual land parcel, y indexes year, and m indexes month. The variable $Monopsony_i$ represents the monopsony power of the land bidder, which is defined as the difference between the transaction price and the reservation price of each land parcel normalized by its reservation price. If the transaction price equals to the reservation price, it suggests that only a few developers bid for the land parcel and the land bidder has a relatively high degree of monopsony power. Therefore, the closer the monopsony measure is to zero, the more bargaining power developers will have in the land auction market. Z_{cy} is a set of city-level time-varying characteristics. A vector of city fixed effects is represented by ϕ_c and a vector of county fixed effects is represented by ρ_d. δ_{ym} is a set of time dummies (year-month fixed effects) and X_i is a set of land parcel controls including land area, distance to CBD, a land quality measured by government, type of auctions, transaction price per square meter, etc. I estimate this equation by OLS, clustering standard errors at the city level to account for potential spatial autocorrelation. The parameter of interest is β, measuring the effect of monopsony power on FAR design.

Table A7 summarizes the results from estimating equation (15) using a sample of land transactions in China between 2007 and 2017. Additional covariates are included into the estimation sequentially. All the estimates for the monopsony power effect are negative and significant, suggesting that developers with higher monopsony power are more likely to negotiate with local governments and purchase land parcels with higher FAR limits.

3.5. Robustness Checks

3.5.1. Spatial Boundary Design

In this sub-section, I conduct two exercises using administrative boundaries to test for the robustness of the main results. One might argue that even the PSM sample from the main analysis is selected based on the observed local characteristics, and there is still a concern of unobserved local features. To mitigate this concern, I apply a spatial boundary design and select land parcels that are geographically close to each other. For instance, Figure B4 presents the land plots within 5 km away from the county boundary within a prefecture, Fuzhou. These land parcels tend to have near-identical neighbourhood and unobserved spatial features and are thus highly comparable. As shown in Figure B5, I select all the land parcels within 5 km away from county boundaries across the country and use these land transactions to perform empirical analysis. The estimation results are reported in Table A8. All the estimated coefficients from columns (1) to (3) are statistically significant and negative, suggesting that the main results are robust after I take into account unobserved local characteristics that might influence both local budgetary revenue and FAR design.

This paper also applies a boundary discontinuity design to visualize the main findings from section 3.3. I select land plots within 10 km away from the boundaries of nine major cities in China. Figure B6 presents the FAR limits for land plots at different distances from the boundaries. Positive distances correspond to locations inside the major city, and negative distances correspond to locations outside the major city. Lines in these figures represent fitted values estimated separately on each side of the boundary. Figure B6 shows clear discontinuities of FAR limits in most major cities at the prefecture boundaries, and I interpret these discontinuities as supportive evidence that higher local budgetary revenues cause lower FAR design in major cities.

3.5.2. Placebo Test

Next, I mitigate the concern that the influence of the TCID policy on FAR design is spuriously documented outside the treatment periods, meaning that the estimated impact might be driven by the pre-existing and unobserved local trends during the process of urbanization, and these pre-trends might not be fully controlled for even after I conduct the propensity score matching. To mitigate the concern of the spurious treatment effect, I generate 1,000 random placebo treatment dates which are 1 to 3 years prior to the real treatment date. For instance, if a city implements the administrative adjustment in 2012, the randomly generated date will be between 2009 and 2011. I then estimate a specification similar to equation (11) using the randomly generated treatment date and the PSM sample. The cumulative probability and the kernel density of the estimated effect from 1,000 different placebo regressions are plotted in Figure B7. The vertical line represents the estimated impact of the TCID policy on FAR design from column (6) of Table A1. Only 16 estimates from these 1,000 placebo regressions are more negative than the estimated treatment effect, increasing the confidence that earlier findings are not spuriously driven by the pre-existing local trends during the urbanization process.

3.5.3. County-Level Estimation Results

To test for the robustness of the main results, this paper aggregates land plot characteristics at the county level by taking the mean value of land plots within each county in each year. This paper then estimates the impacts of budgetary revenue on FAR design using this aggregated county-level panel, and the results are reported in Table A9. In line with the baseline estimates, I find a significant and negative impact of budgetary revenue on FAR limit. Both the IV and the OLS estimates are robust and suggest that a one standard deviation increase in local budgetary revenue will decrease FAR limit by between -0.25 and -0.08.

3.5.4. Drop Tier-1 cities and Municipalities

Four Tier-1 cities (Beijing, Shanghai, Guangzhou, and Shenzhen) and two municipalities (Tianjin and Chongqing) are reckoned as the most economically developed cities in China. To avoid potential bias caused by the unobserved features in these six superstar cities, this paper conducts a robustness check by estimating a sample excluding land transactions in these cities. The results are reported in Table A10. In line with the baseline findings, the estimated coefficients are all negative and statistically significant at 1% level, suggesting that the main results are robust after I mitigate potential bias introduced by superstar cities.

3.5.5. Before the Boom of Local Government Debt

This paper also conducts a robustness check to address the concern of local government debt. After the 2008 financial crisis, the central government in China launched a fiscal stimulus program named the 'four trillion stimulus package' to boost economy. Followed by this stimulus program, local governments in China have increasingly issued debts to finance infrastructure investments, and most of these debts are guaranteed by future land sale revenue. Due to the fiscal pressure of repaying local debts, governments might set relatively high FAR limits to acquire more land sale revenue. As shown in Figure B8, the large-scale issue of local government debts started in 2014. This paper thus estimates a subsample of land transactions between 2005 and 2013 to test for the robustness of the main findings. During this period, local government debt is not likely to have a major impact on FAR design. Table A11 reports the results and shows that all estimates are negative and statistically significant. In line with the baseline estimation, Table A11 suggests that budgetary revenue has a negative effect on FAR design before the boom of local government debt in China.

3.5.6. Prefecture-Level Budgetary Revenue

Lastly, this paper uses the prefecture-level budgetary revenue measure and re-estimates a specification similar to equation (9) to test for the robustness of the main findings. The results are reported in Table A12. All the estimated coefficients are still negative and statistically significant, suggesting that the impact of local budgetary revenue on FAR design is robust after I measure local fiscal capacity at a higher administrative level.

4. Conclusion

This paper studies the determinants of FAR limits in China, a developing country that experiences a rapid process of urbanization in the past decades. I propose a spatial equilibrium framework showing that local governments trade-off between the benefits and the costs of high FAR design. Cities with sufficient budgetary revenue opt to set relatively low FAR limits to reduce the negative externalities caused by construction density. Exploiting a comprehensive dataset of land transactions and a PSM-IV strategy, I find that a one standard deviation increase in local budgetary revenue decreases FAR limits by 0.6.

This paper also provides suggestive evidence on the consequences of the FAR designation process and the 'Land Finance Model' in China. On one hand, Tier-1 cities and cities along the southeast coast have sufficient local budgetary revenue and design relatively low FAR limits, which reduce housing supply and push up housing prices. As Figures B1 and B2 show, cities along the southeast coast have higher housing prices and face more severe housing affordability problems compared with other cities in the country. Restrictive land use regulations also lead to the wealth inequality between homeowners and young first-time buyers within these cities. On the other hand, some local governments in the less developed western and middle regions cannot collect sufficient budgetary revenue from local taxes and choose to design higher FAR limits to acquire more land sale revenue. These cities are not experiencing economic prosperity as Beijing and Shanghai and thus cannot attract superstar firms and high-skilled labours. As a result, many high-rise buildings in these cities are constructed and then left vacant. Despite the staggeringly high housing prices in Tier-1 cities, properties in some lower-Tier cities are sold for only 300 RMB/m^2, which is around 33 GBP/m^2 (Xinhuanet, 2019).[51] The regional inequality between the 'under-occupied

[51] Based on the currency exchange rate in September 2021.

cities' and the 'unaffordable cites' is largely driven by the fact that land sale serves as a major source of fiscal revenue for many local governments. While this paper studies the determinants of FAR design, future research can explore how Chinese cities can improve the current 'Land Finance Model' and develop in a sustainable way.

Tables

Table 1: Descriptive Statistics

	Observations	Mean	SD	Max	Min
Panel A: Land parcel characteristics					
FAR upper limit	202816	2.8	1.4	8	1
Transaction price (10,000 RMB)	202816	5300	19977.1	1020000	0.0001
Distance to CBD (km)	202816	42.2	27.6	100	0.0001
Land area (10,000 m^2)	202816	2.5	94.7	42559	0.00004
Auction type 1 (zhao)	202816	0.01	0.1	1	0
Auction type 2 (pai)	202816	0.2	0.4	1	0
Auction type 3 (gua)	202816	0.5	0.5	1	0
Auction type 4 (negotiation)	202816	0.3	0.4	1	0
Auction type 5 (huabo)	202816	0.01	0.1	1	0
Land quality	202816	4.6	4.4	18	0
Panel B: Prefectural-level characteristics					
Budgetary revenue (1 million RMB)	3027	17425.8	41324.6	664226.4	208.2
Population (1 million)	3027	4.5	3.1	33.9	0.2
% Employment in the agricultural sector	3027	2.9	6.6	74	0
% Employment in the tertiary sector	3027	52.5	12.9	94.8	9.9
Average salary (RMB)	3027	40338.3	17578.9	320626.3	4958
Number of universities	3027	8.5	14.4	92	1
Number of hospitals	3027	212.5	188	3052	5
Panel C: County-level characteristics					
Budgetary revenue (1 million RMB)	15380	1312.6	2613.6	67298.4	11.2
TCID policy treatment dummy	1823	0.1	0.3	1	0

Table 2: The Effect of Budgetary Revenue on FAR

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	IV	IV	IV
Budgetary revenue	-0.0756***	-0.0834***	-0.0772***	-0.4296***	-0.5248***	-0.6013***
	(0.0120)	(0.0137)	(0.0136)	(0.0909)	(0.1228)	(0.1601)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
Grid FEs	No	Yes	Yes	No	Yes	Yes
City controls [2)]	No	No	Yes	No	No	Yes
Number of counties	1804	1791	1791	1804	1791	1791
Number of grids		16182	16182		16182	16182
Time period			2005-2017			
N	202797	195070	195070	202797	195070	195070
R^2	0.4024	0.5336	0.5340			
Kleibergen-Paap rk Wald F-statistic				90.08	67.24	39.77

Notes: [1)] Land parcel controls include land area, distance to CBD, type of auction, land quality, longitude and latitude of the land parcel. [2)] City controls include population, average salary, local industry composition, number of universities and number of hospitals. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 3: First Stage Results

Specifications	(1)	(2)	(3)
Dependent variable	Budgetary revenue	Budgetary revenue	Budgetary revenue
TCID × Post	0.4968***	0.4066***	0.3197***
	(0.0523)	(0.0496)	(0.0507)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
County FEs	Yes	Yes	Yes
Grid FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
N	202797	195070	195070
R^2	0.9040	0.9347	0.9369

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 4: Propensity Score Matching Results

	Before Matching			After Matching		
Variables	Treated cities	Control cities	P value	Treated cities	Control cities	P value
Variables used in PSM						
Population (10,000)	150	107	0.001	150	160	0.72
Pop. growth rate (%)	3.2	1.7	0.18	3.2	4	0.72
Budgetary revenue (1 million RMB)	9374	4220	0	9374	7754	0.53
GDP per capita (RMB)	51710	41307	0.01	51710	46625	0.45
% Employment (secondary industry)	46.4	45.7	0.74	46.4	45.4	0.69
% Employment (tertiary industry)	53	52.5	0.82	53	54	0.67
Variables not used in PSM						
Population density per km^2	463	399	0.13	463	568	0.16
Average salary (RMB)	32230	30417	0.04	32230	32194	0.98

Table 5: The Effect of Budgetary Revenue on FAR (PSM Sample)

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	IV	IV	IV
Budgetary revenue	-0.0142	-0.0418*	-0.0232	-0.4321**	-0.6085***	-0.6153***
	(0.0229)	(0.0253)	(0.0260)	(0.2076)	(0.2264)	(0.2363)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
Grid FEs	No	Yes	Yes	No	Yes	Yes
City controls [2)]	No	No	Yes	No	No	Yes
Number of counties	216	216	216	216	216	216
Number of grids		2213	2213		2213	2213
Time period	2005-2017					
N	22016	21260	21260	22016	21260	21260
R^2	0.3279	0.4652	0.4663			
Kleibergen-Paap rk Wald F-statistic				31.84	26.74	28.11

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 6: First Stage Results (PSM Sample)

Specifications	(1)	(2)	(3)
Dependent variable	Budgetary revenue	Budgetary revenue	Budgetary revenue
TCID × Post	0.2123***	0.2232***	0.2123***
	(0.0376)	(0.0432)	(0.0400)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
County FEs	Yes	Yes	Yes
Grid FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
N	22016	21260	21260
R^2	0.8321	0.8701	0.8799

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Figures

Fig. 1: Fiscal Revenue and Land Revenue in China

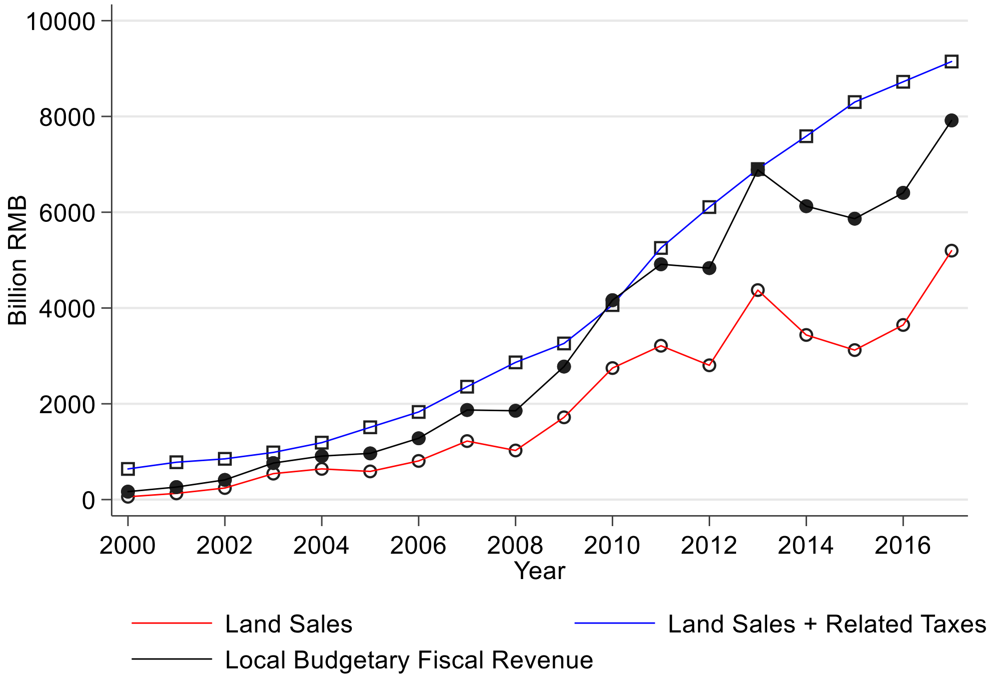


Fig. 2: Construction Density and Air Quality

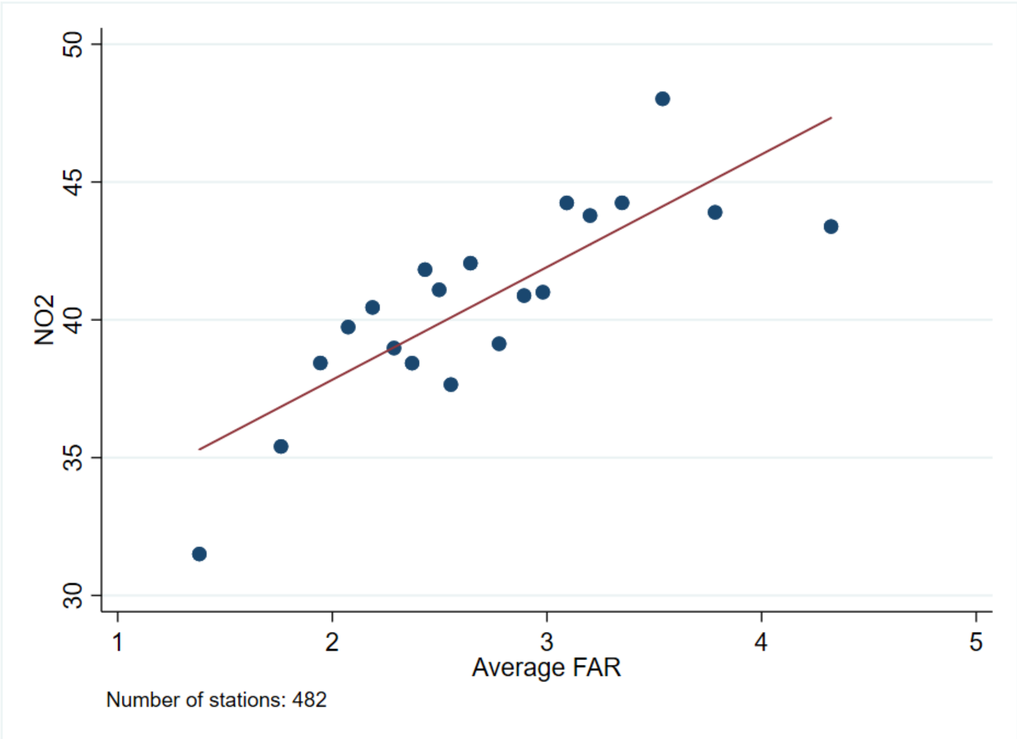


Fig. 3: FAR Limit and Residential Land Price

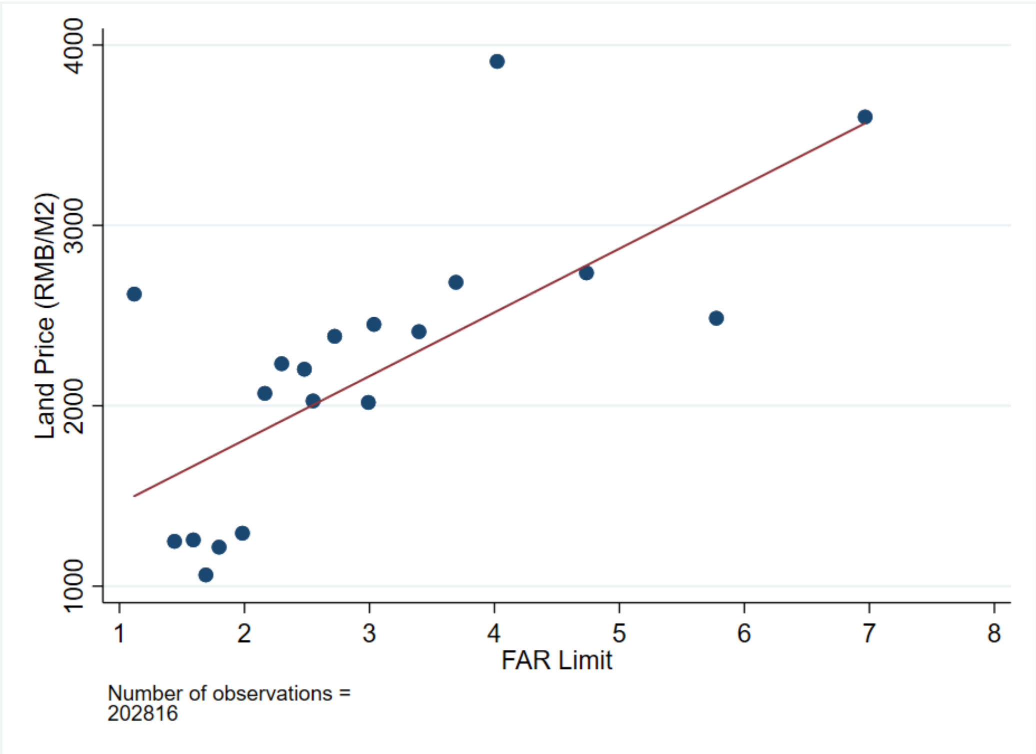


Fig. 4: Prefectural-level Budgetary Revenue in China in 2016

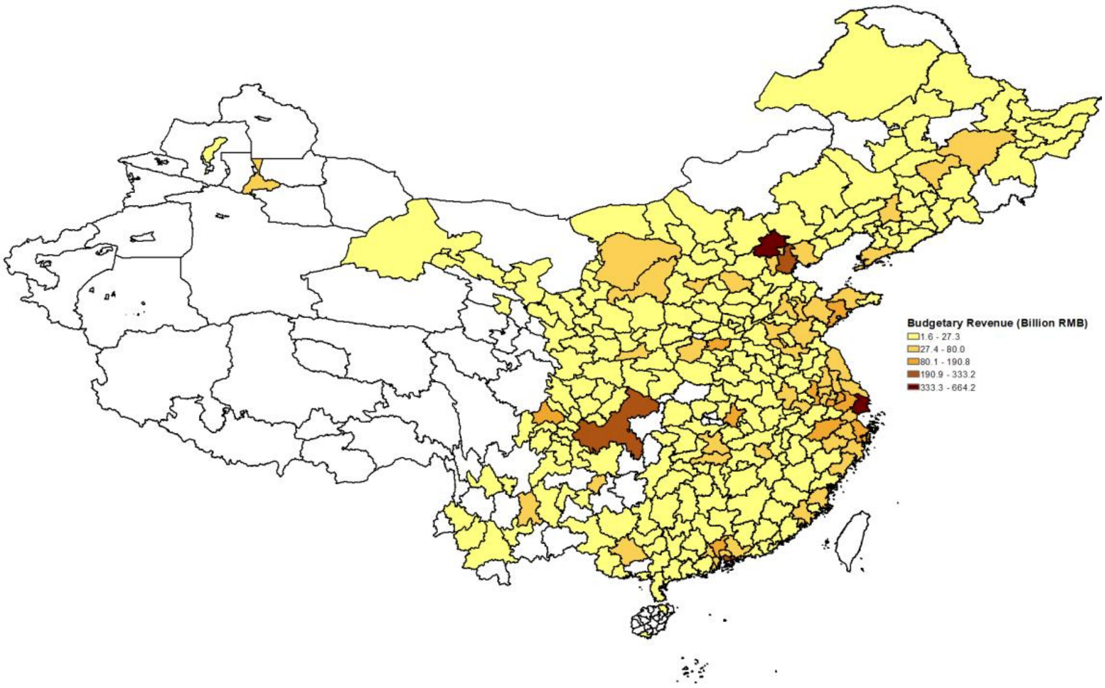


Fig. 5: Weighted Average FAR in China (Residential Use)

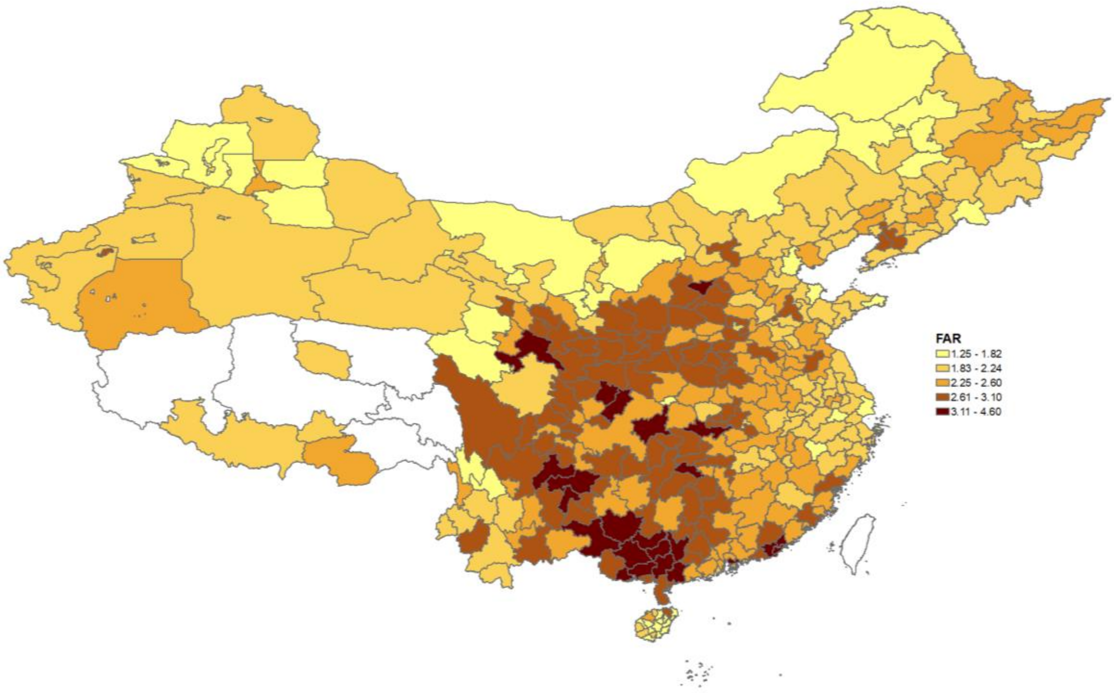


Fig. 6: Budgetary Revenue and FAR Limit

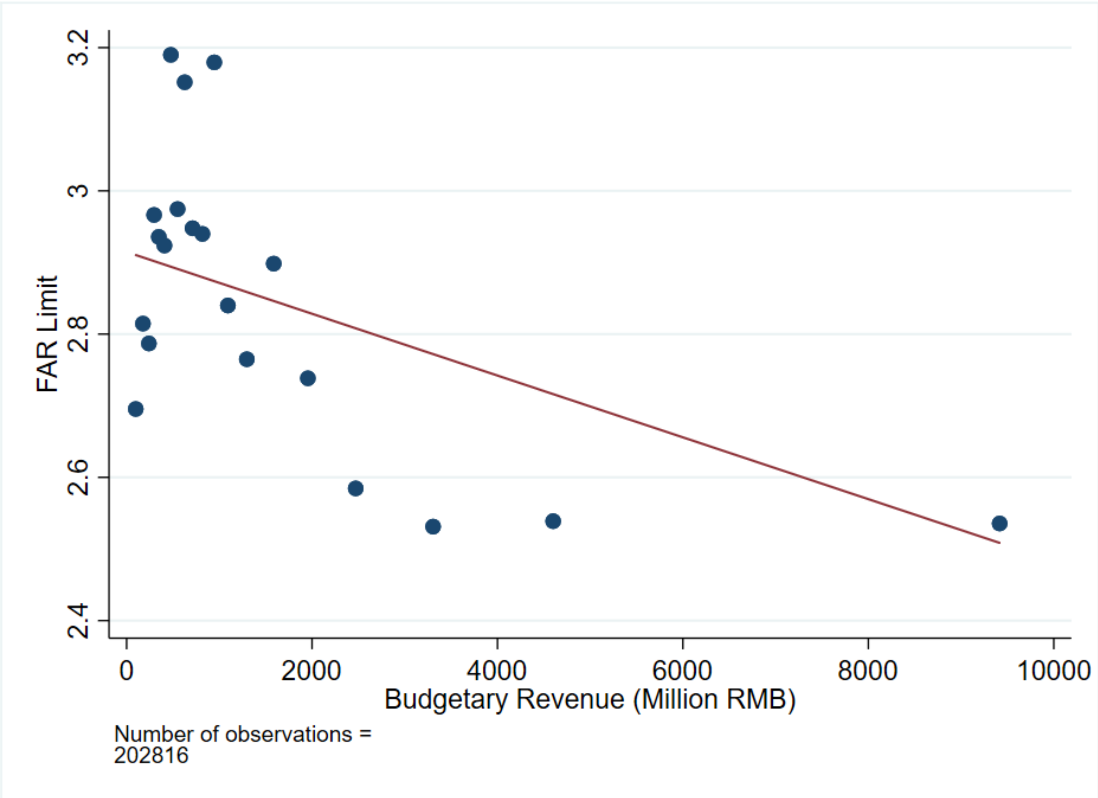


Fig. 7: Local Administrative Division in China

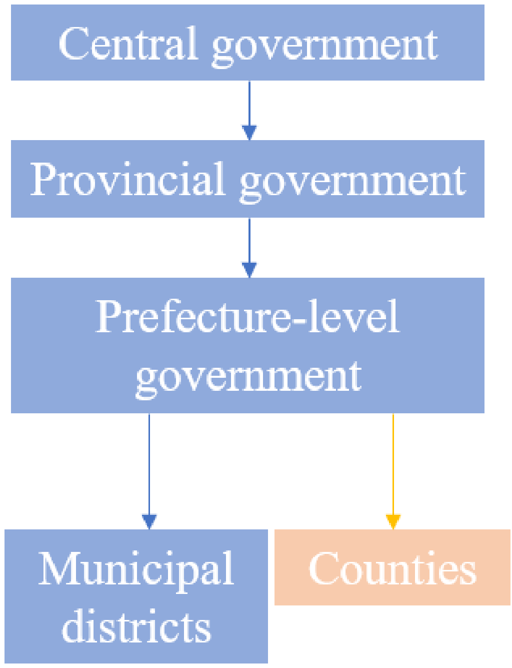


Fig. 8: The Number of Administrative Adjustments

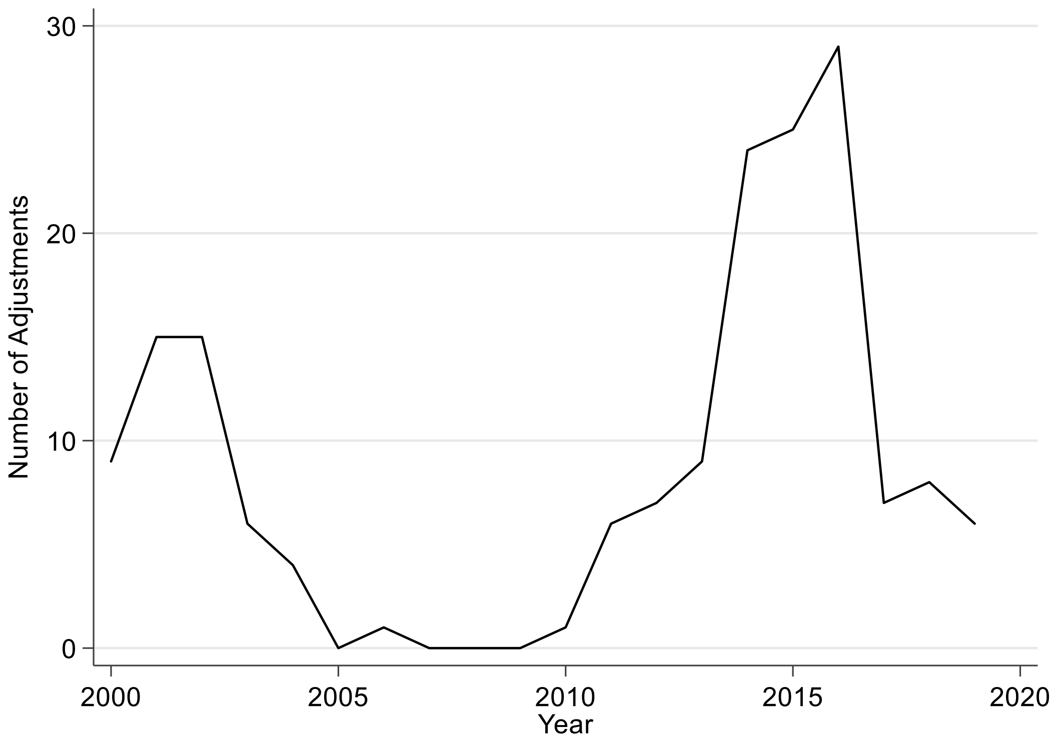


Fig. 9: Cities with the TCID Policy in China

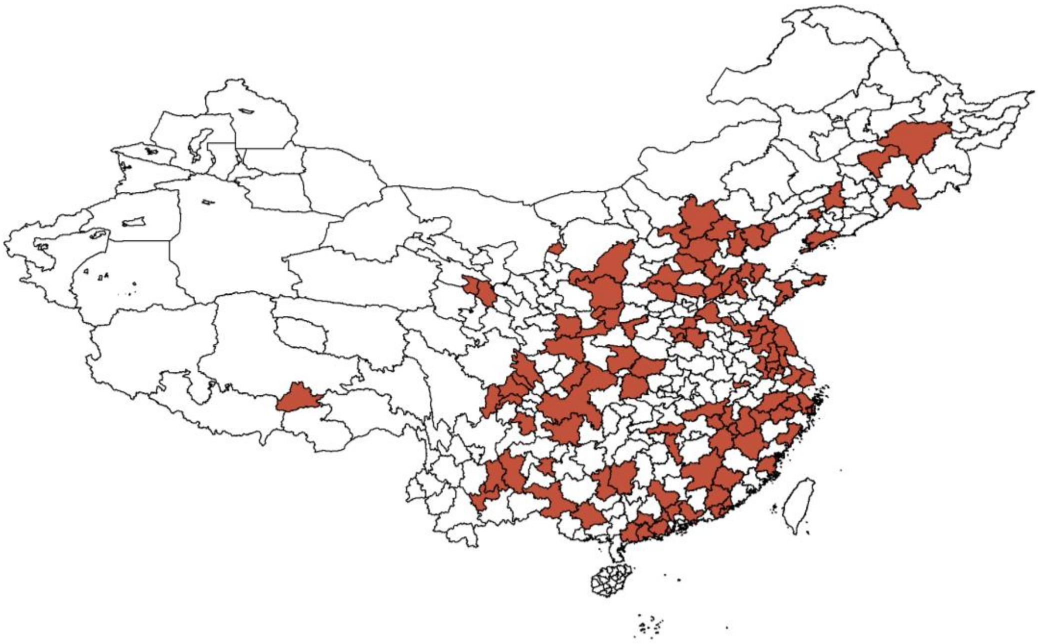


Fig. 10: Turning County into District – Example of Xi'an

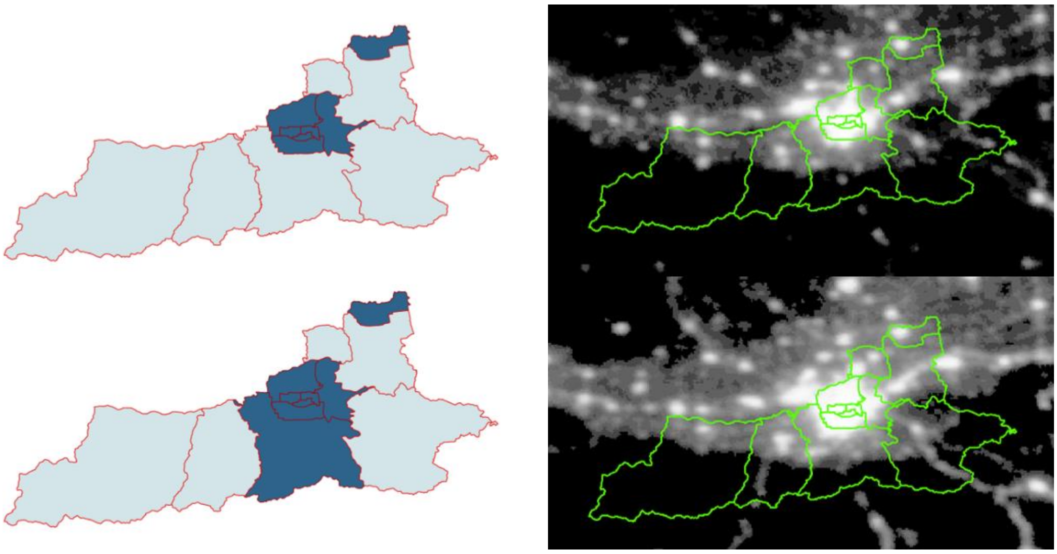


Fig. 11: The Effects of FAR Design

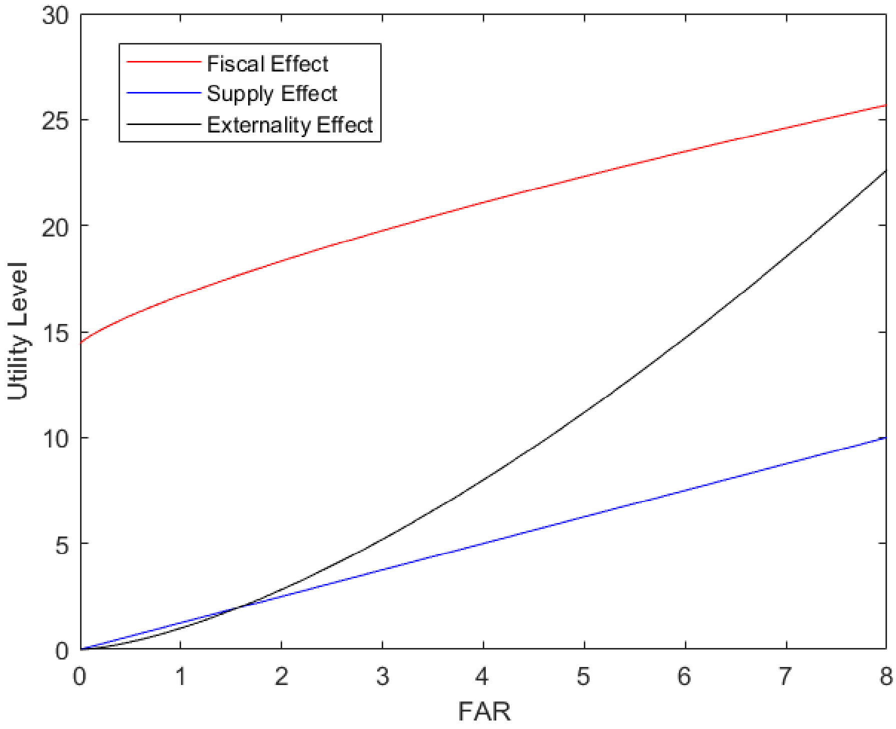


Fig. 12: Geocoded Land Parcels

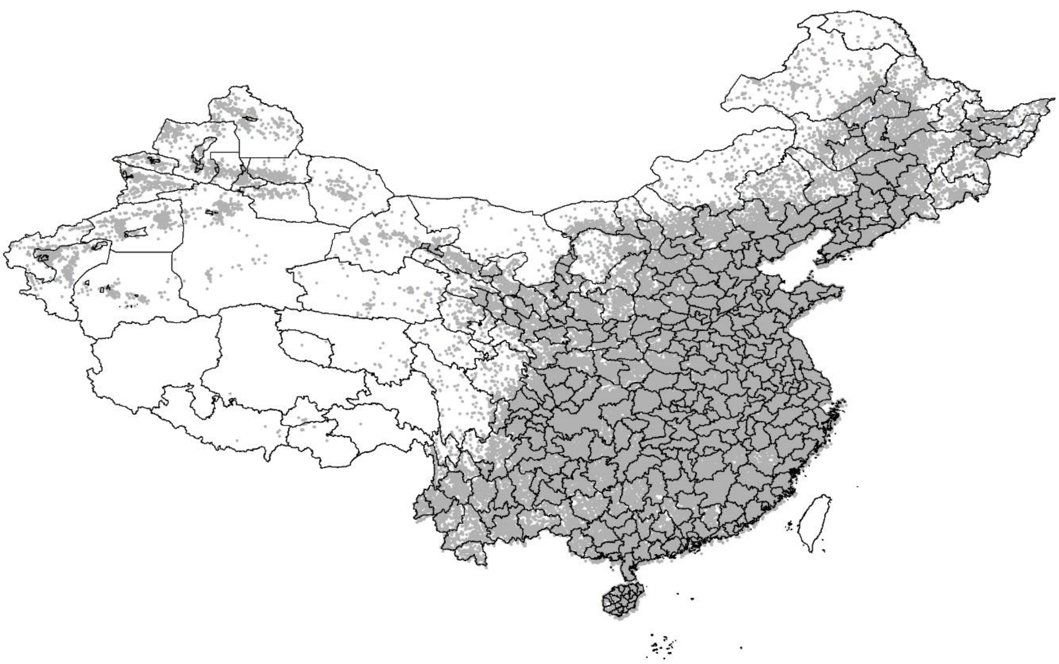


Fig. 13: Land Parcels and 3km × 3km Grids in Beijing

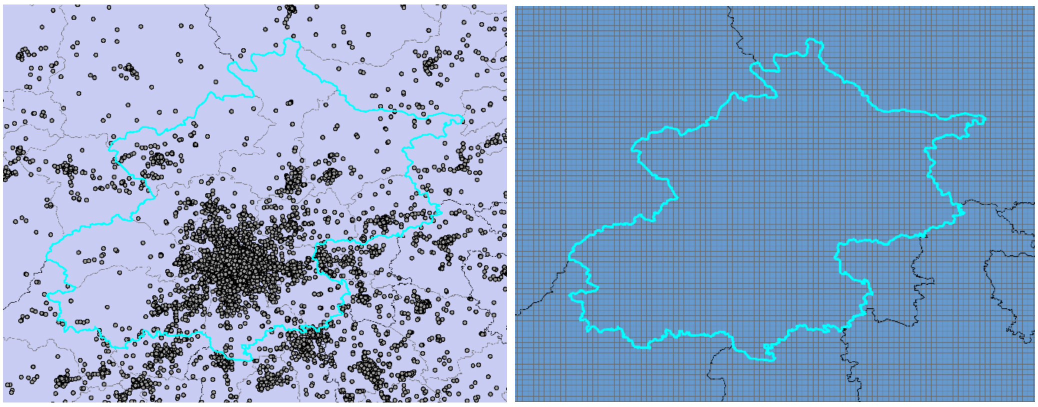


Fig. 14: Histogram of FAR Restriction

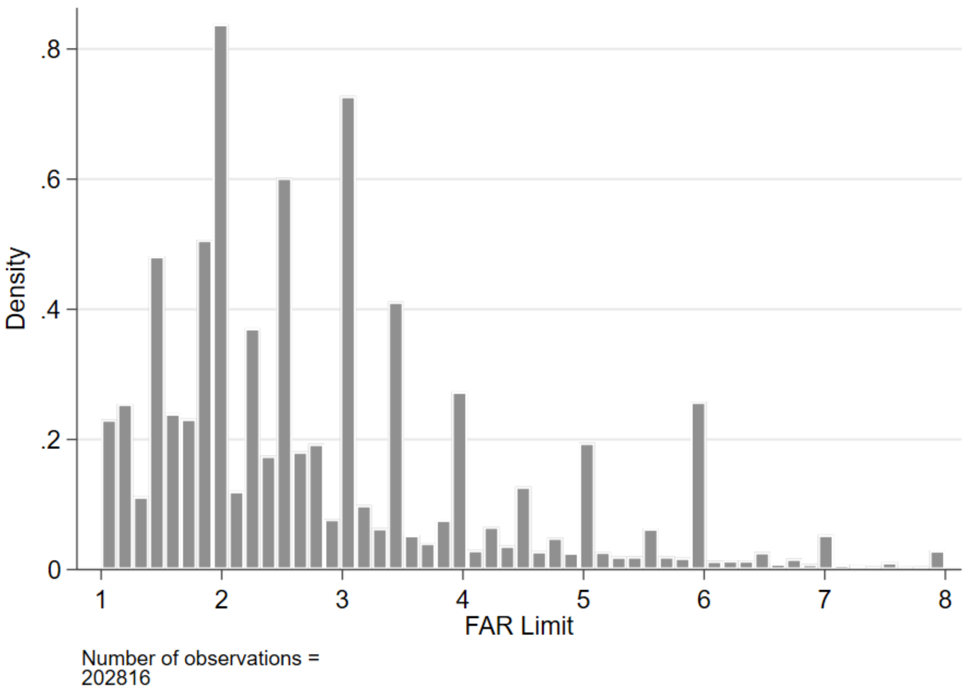


Fig. 15: Average Budgetary Revenue (Before PSM)

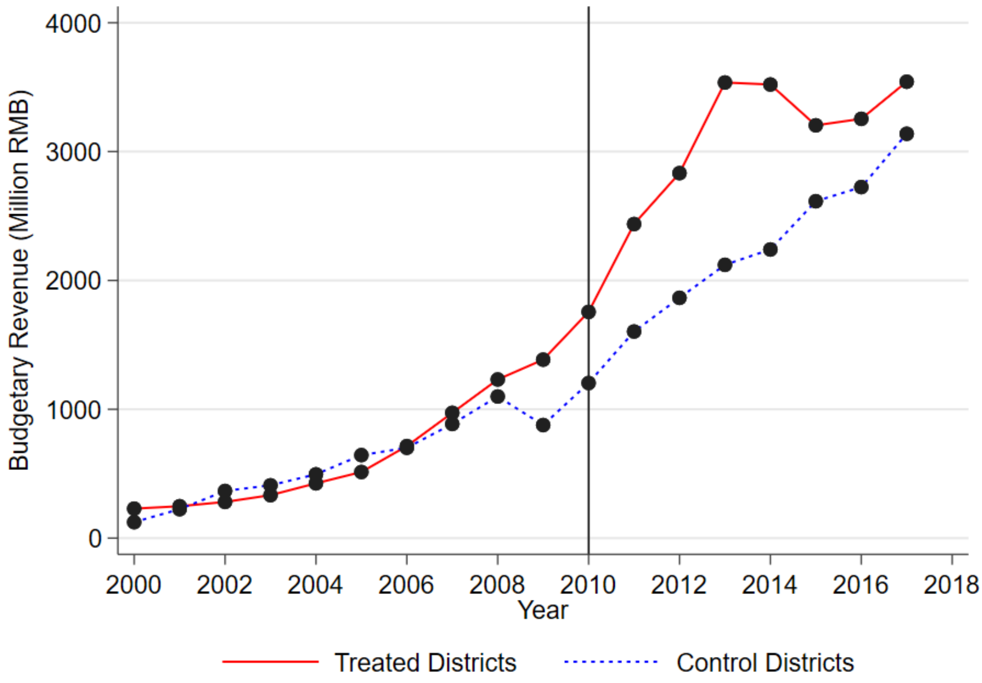


Fig. 16: Average Budgetary Revenue (After PSM)

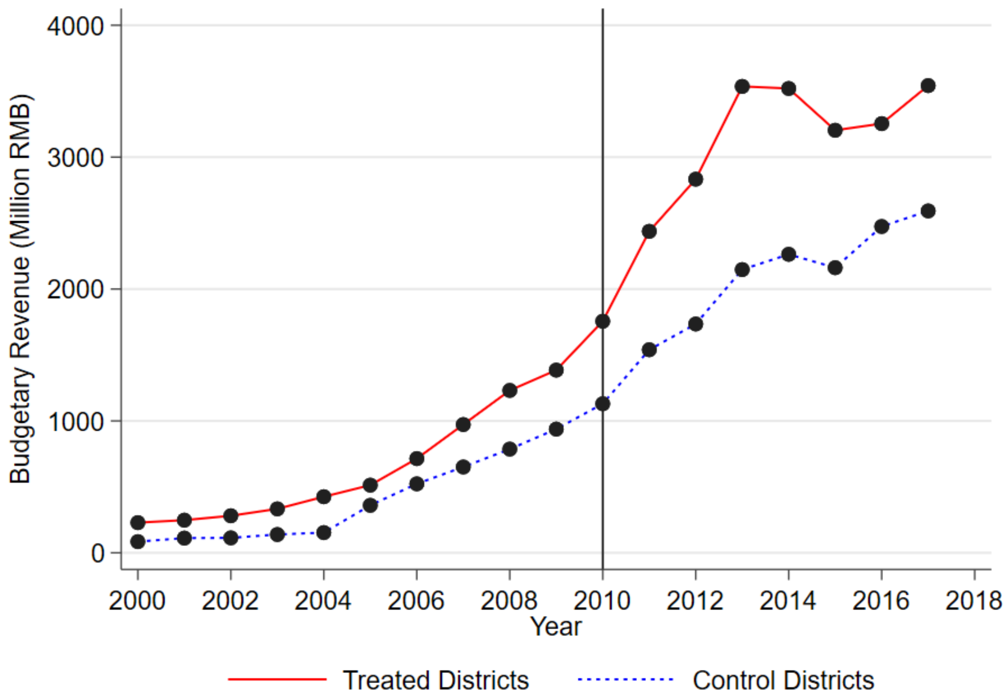


Fig. 17: Annual Average FAR between 2007 and 2018 (PSM Sample)

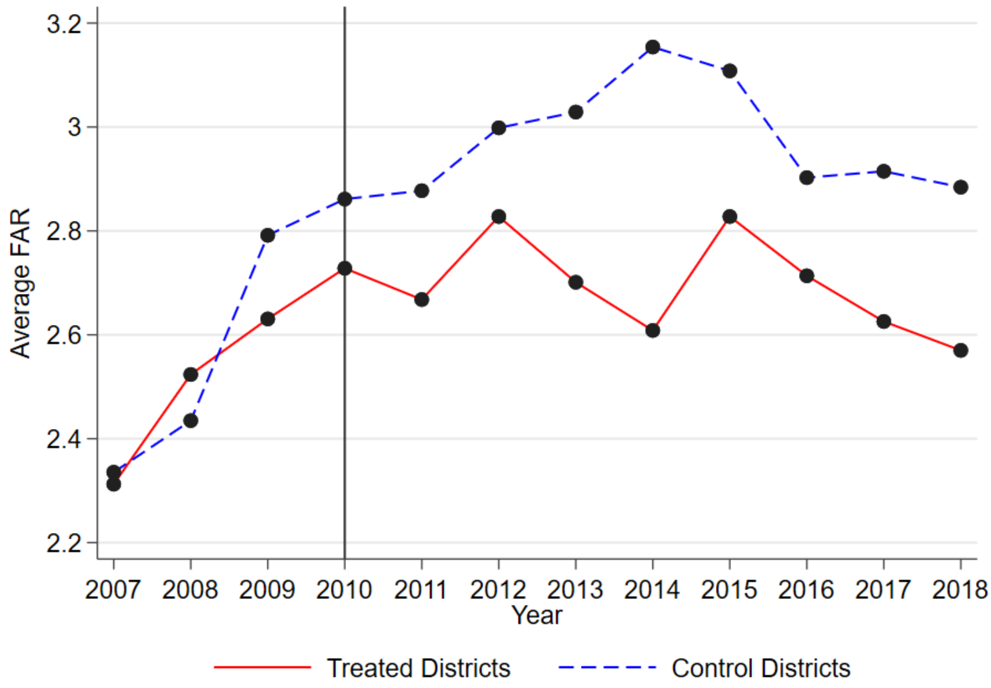
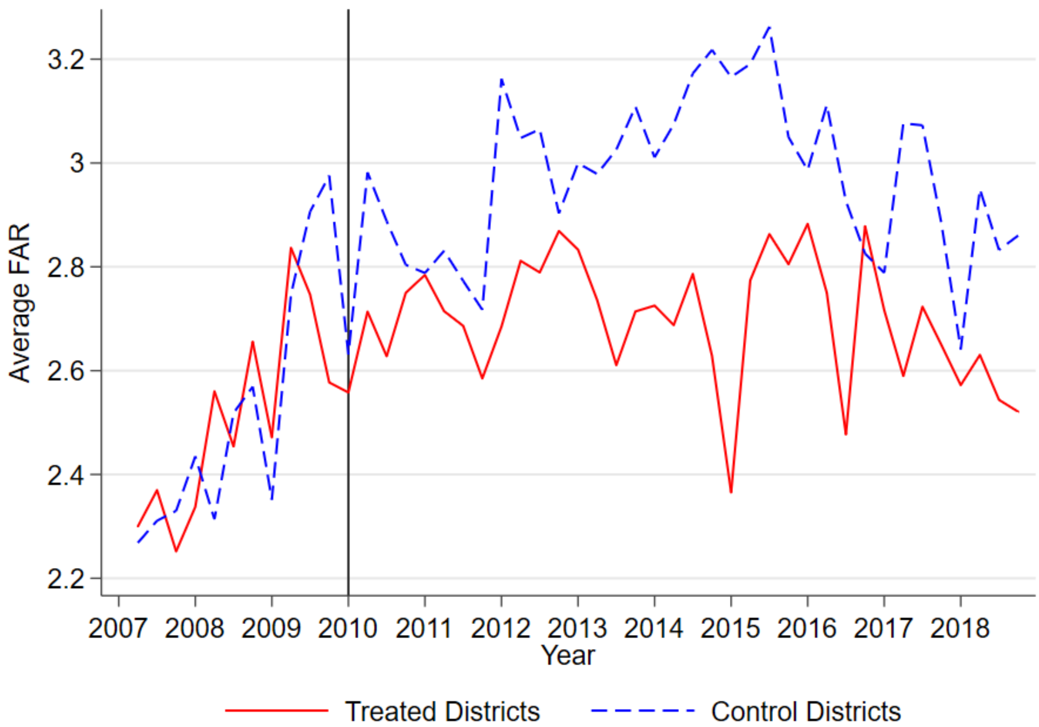


Fig. 18: Quarterly Average FAR between 2007 and 2018 (PSM Sample)



Appendices

Appendix A: Appendix Tables

Table A1: The Effect of Administrative Adjustment on FAR

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	OLS	OLS	OLS
Sample	Baseline Sample			PSM Sample		
TCID × Post	-0.1855***	-0.1967***	-0.1733***	-0.0917**	-0.1358***	-0.1306***
	(0.0591)	(0.0639)	(0.0621)	(0.0431)	(0.0464)	(0.0458)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
Grid FEs	No	Yes	Yes	No	Yes	Yes
City controls [2)]	No	No	Yes	No	No	Yes
Number of counties	1804	1791	1791	216	216	216
Number of grids		16182	16182		2213	2213
Time period	2005-2017			2005-2017		
N	202797	195070	195070	22016	21260	21260
R^2	0.4022	0.5335	0.5339	0.3282	0.4655	0.4667

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A2: Population Density and Negative Externality

Specifications	(1)	(2)	(3)	(4)
	Log(PM10)	Log(PM10)	Log(AQI)	Log(AQI)
Log(population density)	0.2372*	0.2413*	0.1836**	0.1848**
	(0.1331)	(0.1292)	(0.0917)	(0.0893)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
City Controls [1)]	No	Yes	No	Yes
N	821	821	821	821
R^2	0.8670	0.8716	0.7407	0.7476

Notes: [1)] City controls include average salary, the proportions of employment in the agricultural industry and the tertiary industry. Standard errors are clustered at the prefecture level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A3: The Effect of Budgetary Revenue on Housing Prices

Specifications	(1)	(2)	(3)	(4)
Dependent variable	Ln(HP)	Ln(HP)	Ln(HP)	Ln(budgetary revenue)
	OLS	OLS	IV	First Stage
Ln(budgetary revenue)	0.0721***	0.0657***	0.2659**	
	(0.0120)	(0.0120)	(0.1109)	
TCID × Post				0.1913***
				(0.0500)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
City Controls [1)]	No	Yes	Yes	Yes
Number of prefectures	283	283	283	283
N	3459	3459	3459	3459
R^2	0.9551	0.9556		0.9677
Kleibergen-Paap rk Wald F-statistic			14.61	

Notes: [1)] City controls include average salary, GDP per capita, the proportions of employment in the agricultural and tertiary sectors. Standard errors are clustered at the prefecture level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A4: The Effect of Budgetary Revenue on FAR (Non-residential Uses)

Specifications	(1)	(2)	(3)	(4)
	FAR	FAR	FAR	FAR
	commercial use	commercial use	industrial use	industrial use
Budgetary revenue	-0.0168	-0.0099	0.0258***	0.0197***
	(0.0131)	(0.0135)	(0.0050)	(0.0054)
Land parcel controls [1)]	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes
Grid FEs	Yes	Yes	Yes	Yes
City controls [2)]	No	Yes	No	Yes
Number of counties	1806		1781	
Number of grids	13210		16981	
Time period	2005-2017		2005-2017	
N	87645	87645	135393	135393
R^2	0.4933	0.4936	0.7172	0.7179

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A5: The Elasticity of Land Price with Respect to FAR

Specifications	(1)	(3)	(4)
Dependent variable	Ln(land price) residential use	Ln(land price) commercial use	Ln(land price) industrial use
Ln(FAR)	0.4281***	0.2856***	0.0717***
	(0.0325)	(0.0094)	(0.0087)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
County FEs	Yes	Yes	Yes
Grid FEs	Yes	Yes	Yes
City controls [2)]	Yes	Yes	Yes
N	195070	87645	135393
R^2	0.7125	0.6989	0.7791

Notes: [1)] See Table 2. [2)] City controls include all the city controls as indicated in Table 2 as well as local budgetary revenue. Standard errors are clustered at the grid level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A6: The Effect of Transfer Payment on FAR

Specifications	(1)	(2)	(3)
	FAR	FAR	FAR
Transfer payment	-0.1502***	-0.1318***	-0.0919***
	(0.0260)	(0.0269)	(0.0288)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
City FEs	Yes	Yes	Yes
Grid FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
Number of prefectures	263	263	263
Number of grids		20044	20044
N	246143	236947	236947
R^2	0.2677	0.5250	0.5253

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The transfer payment variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A7: The Effect of Monopsony Power on FAR

Specifications	(1)	(2)	(3)
Monopsony power	-0.1284***	-0.1332***	-0.1332***
	(0.0410)	(0.0386)	(0.0386)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
City FEs	Yes	Yes	Yes
County FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
N	97842	97617	97617
R^2	0.2461	0.3871	0.3871

Notes: [1)] Land parcel controls include land area, distance to CBD, land price per square meter, type of auction. [2)] See Table 2. Standard errors are clustered at the prefecture level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A8: The Effect of Budgetary Revenue on FAR (Spatial Boundary Design)

Specifications	(1)	(2)	(3)
	OLS	OLS	OLS
Budgetary revenue	-0.0733*	-0.1029**	-0.1088**
	(0.0377)	(0.0469)	(0.0491)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
County FEs	Yes	Yes	Yes
Grid FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
Number of counties	666	622	622
Number of grids		1837	1837
Time period		2005-2017	
N	18769	17772	17772
R^2	0.4592	0.5551	0.5565

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A9: County-Level Estimation Results

Specifications	(1) OLS	(2) OLS	(3) IV	(4) IV	(5) First stage	(6) First stage
Budgetary revenue	-0.0922***	-0.0808***	-0.2474***	-0.2268**		
	(0.0146)	(0.0140)	(0.0781)	(0.0888)		
TCID × Post					0.7475***	0.6329***
					(0.1104)	(0.1033)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
City controls [2)]	No	Yes	No	Yes	No	Yes
Number of counties	1789					
Time period	2005-2017					
N	15345	15345	15345	15345	15345	15345
R^2	0.5702	0.5712			0.8908	0.8951
Kleibergen-Paap rk Wald F-statistic			45.82	37.55		

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the county level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A10: Robustness Check: Drop Tier-1 Cities and Municipalities

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	IV	IV	IV
Budgetary revenue	-0.0841***	-0.0918***	-0.0899***	-0.4270***	-0.5370***	-0.5691***
	(0.0151)	(0.0153)	(0.0155)	(0.1283)	(0.1526)	(0.1704)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
Grid FEs	No	Yes	Yes	No	Yes	Yes
City controls [2)]	No	No	Yes	No	No	Yes
Number of counties	1747	1740	1740	1747	1740	1740
Number of grids		15803	15803		15803	15803
Time period	2005-2017					
N	200039	192559	192559	200039	192559	200039
R^2	0.4037	0.5339	0.5342			
Kleibergen-Paap rk Wald F-statistic				72.61	47.85	38.42

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A11: Robustness Check (Sample before 2014)

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	IV	IV	IV
Budgetary revenue	-0.0345**	-0.0400**	-0.0339*	-0.7294**	-0.8205**	-1.2887*
	(0.0165)	(0.0195)	(0.0191)	(0.2981)	(0.3208)	(0.6819)
Land parcel controls [1)]	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes	Yes	Yes	Yes
County FEs	Yes	Yes	Yes	Yes	Yes	Yes
Grid FEs	No	Yes	Yes	No	Yes	Yes
City controls [2)]	No	No	Yes	No	No	Yes
Number of counties	1722	1696	1696	1722	1696	1696
Number of grids		11975	11975		11975	11975
Time period	2005-2013					
N	128078	121524	121524	128078	121524	121524
R^2	0.4443	0.5709	0.5715			
Kleibergen-Paap rk Wald F-statistic				33.95	22.52	7.08

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A12: Robustness Check (The Effect of Prefecture Level Budgetary Revenue on FAR)

Specifications	(1)	(2)	(3)
	OLS	OLS	OLS
Budgetary revenue	-0.0724***	-0.0726***	-0.0542***
	(0.0124)	(0.0131)	(0.0129)
Land parcel controls [1)]	Yes	Yes	Yes
Year-month FEs	Yes	Yes	Yes
City FEs	Yes	Yes	Yes
Grid FEs	No	Yes	Yes
City controls [2)]	No	No	Yes
Number of prefectures	281	281	281
Number of grids		20823	20823
Time period		2005-2017	
N	253614	243990	243990
R^2	0.2706	0.5266	0.5269

Notes: [1)] See Table 2. [2)] See Table 2. Standard errors are clustered at the grid level. The budgetary revenue variable is standardized. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Appendix B: Appendix Figures

Fig. B1: Housing Price in China in 2017

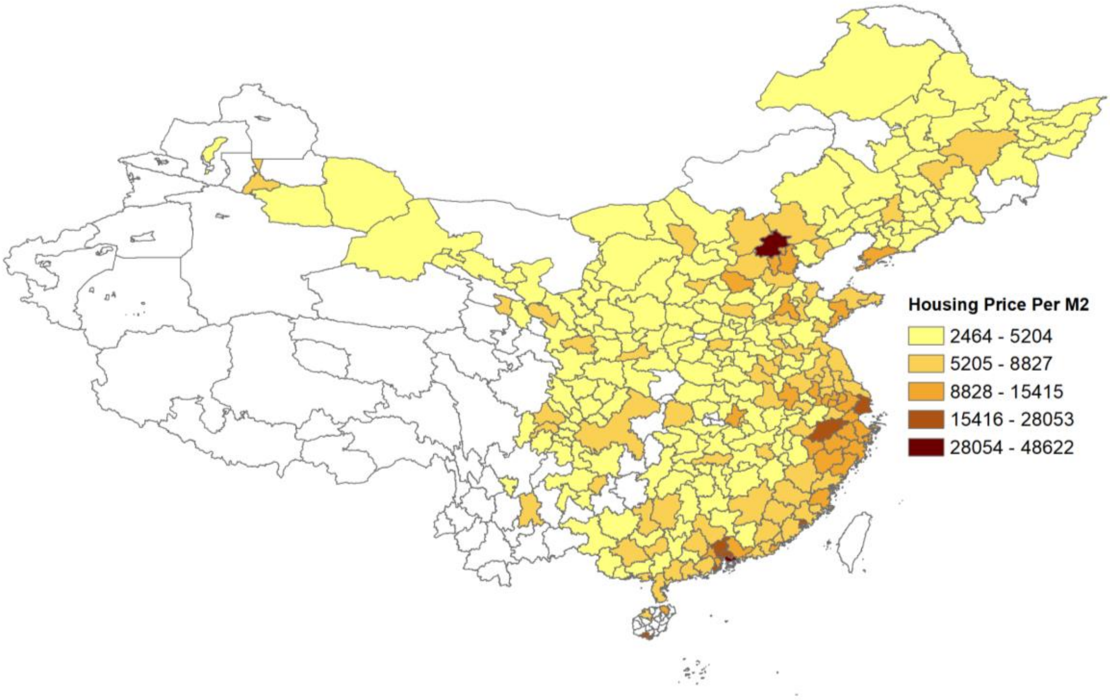


Fig. B2: Housing Affordability in China in 2017

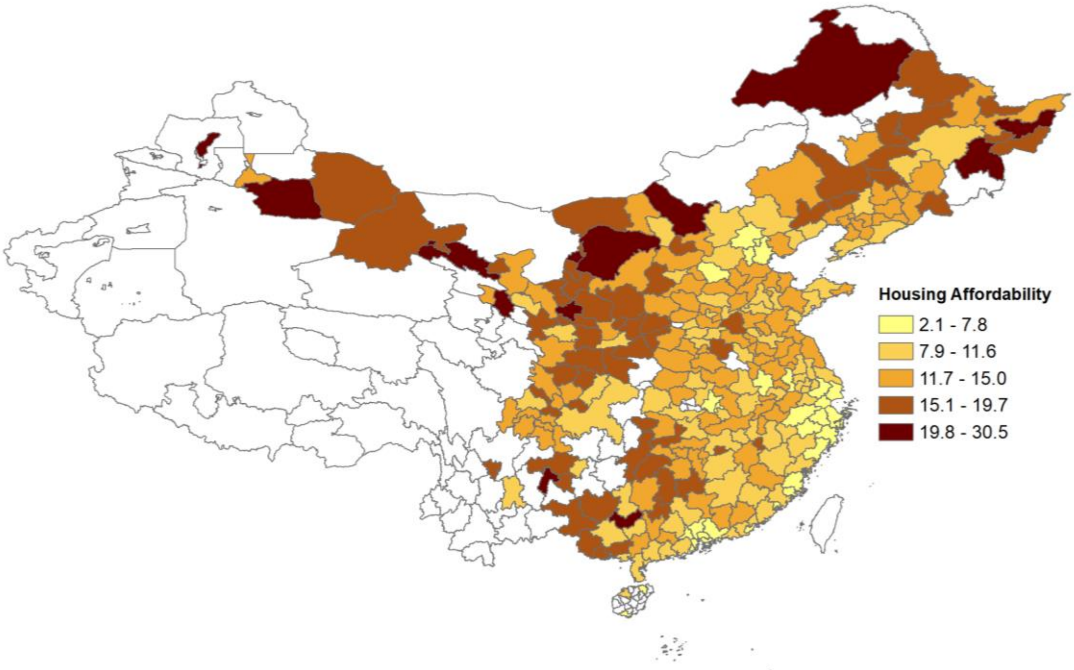


Fig. B3: Weighted Average FAR in China (Commercial and Industrial Uses)

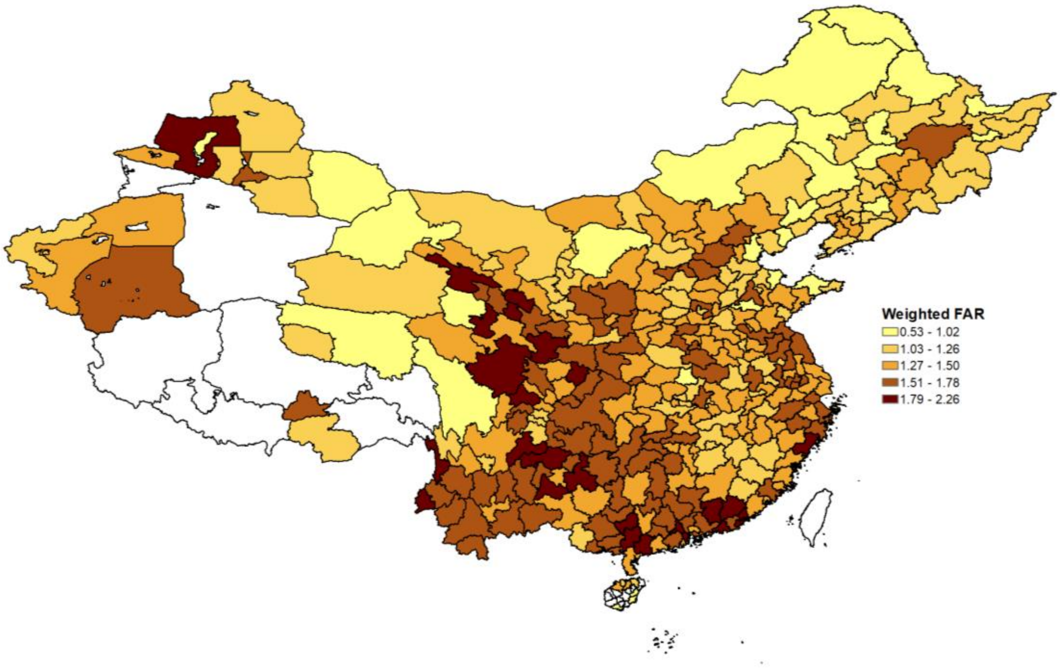


Fig. B4: Land Plots Close to the County Boundary in Fuzhou (Jiangxi Province)

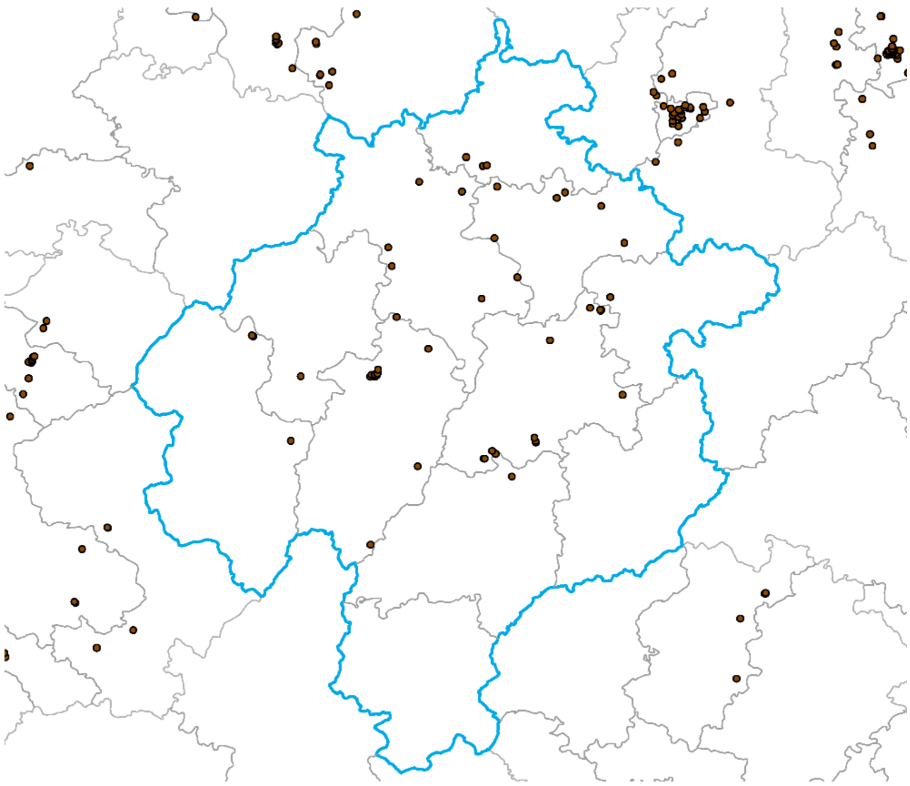
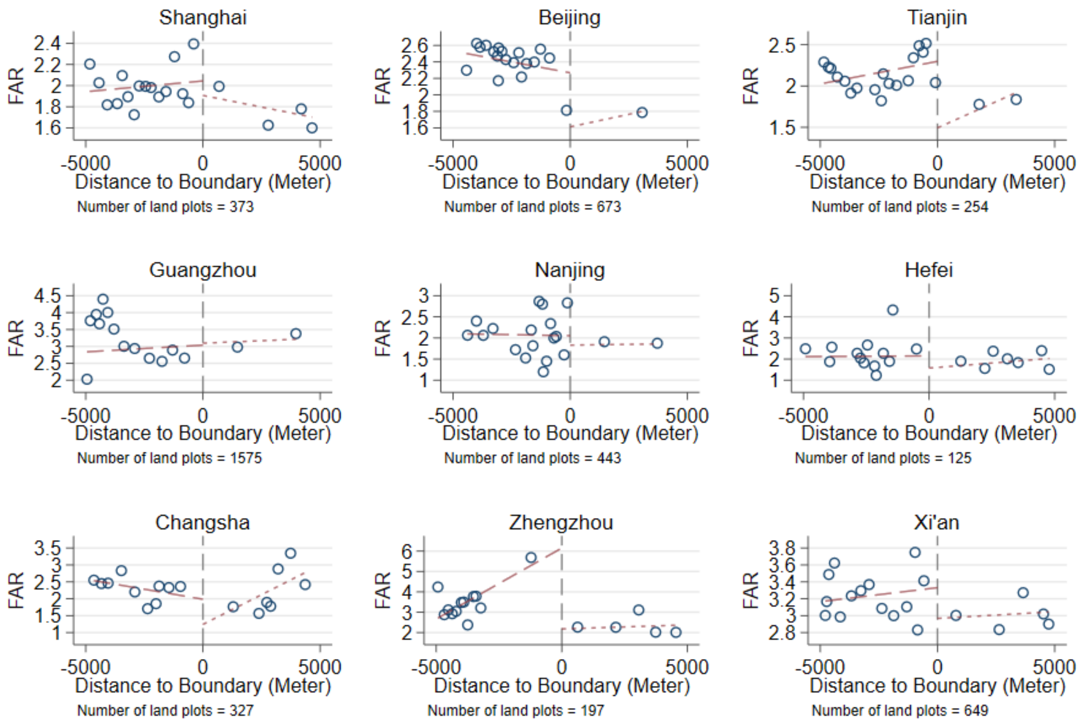


Fig. B5: Land Plots Close to the County Boundary in China

Fig. B6: Spatial Discontinuity of FAR at Prefecture Boundaries



Note: positive distance represents land plots within richer cities.

Fig. B7: Placebo Test

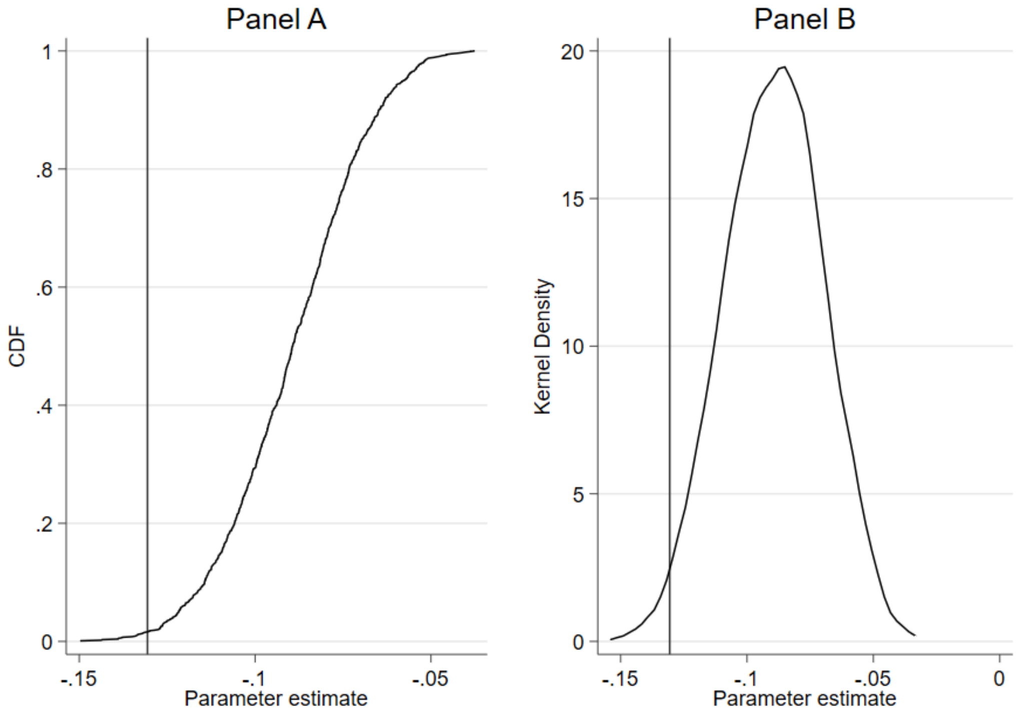
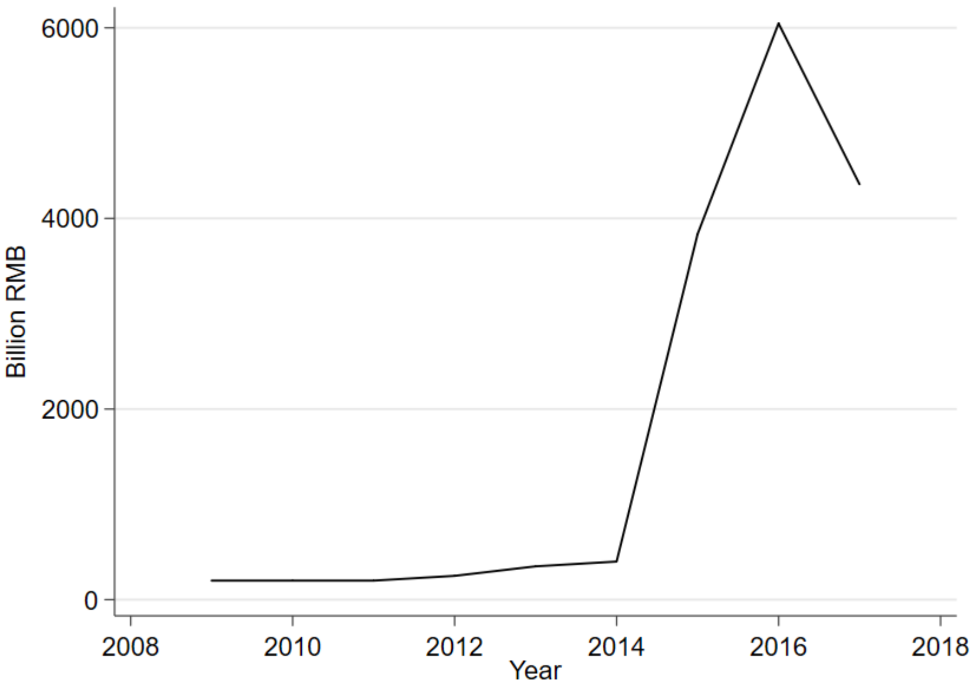


Fig. B8: Local Government Debt in China



Note: Estimates from Suning Institute of Finance

Appendix C: Theoretical Appendix

To prove $\frac{\partial f^*}{\partial B} < 0$

Equation (6) determines the distribution of population across cities under the spatial equilibrium:

$$\bar{U} = \frac{\left(\frac{1-\beta}{\beta}\right)^{1-\beta} (NSf + d)}{L} + \left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma} - e(f)$$

Let ψ denote $\left(\frac{1-\beta}{\beta}\right)^{1-\beta}$. Local government design FAR f to maximize population. The first order condition is:

$$-\frac{\psi(NSf + d)}{L^2}\frac{\partial L}{\partial f} + \frac{\psi NS}{L} + A_G \gamma NS \left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma-1} \frac{\partial r}{\partial f} - \gamma \frac{A_G NSr + A_G B}{L^2}\left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma-1} \frac{\partial L}{\partial f} - e'(f) = 0$$

At the optimal point, $\frac{\partial L}{\partial f} = 0$, therefore:

$$\frac{\psi NS}{L} + A_G \gamma NS \left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma-1} \frac{\partial r}{\partial f} - e'(f) = 0 \qquad (16)$$

$$\left(\frac{A_G NSr + A_G B}{L}\right)^{\gamma-1} = \frac{e'(f) - \frac{\psi NS}{L}}{A_G NS \frac{\partial r}{\partial f}}$$

$$B = \frac{L}{A_G}\left(\frac{e'(f) - \frac{\psi NS}{L}}{A_G NS \frac{\partial r}{\partial f}}\right)^{\frac{1}{\gamma-1}} - NSr$$

The derivative of B with respect to FAR f is thus:

$$\frac{\partial B}{\partial f} = \frac{L}{A_G(\gamma-1)}\left(\frac{e'(f) - \frac{\psi NS}{L}}{A_G NS \frac{\partial r}{\partial f}}\right)^{\frac{2-\gamma}{\gamma-1}} \left(\frac{e''(f)}{A_G NS \frac{\partial r}{\partial f}} - \frac{\partial^2 r}{\partial^2 f} \times \frac{e'(f) - \frac{\psi NS}{L}}{\left(A_G NS \frac{\partial r}{\partial f}\right)^2}\right) - NS\frac{\partial r}{\partial f} \qquad (17)$$

$$\frac{\partial r}{\partial f} > 0, \frac{\partial^2 r}{\partial^2 f} < 0$$

$$e'(f) > 0, e''(f) > 0$$

$$\gamma - 1 < 0$$

Plug equation (16) into equation (17):

$$\frac{\partial B}{\partial f} = \frac{L}{A_G(\gamma-1)}\left(\frac{A_G\gamma NS\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}\frac{\partial r}{\partial f}}{A_GNS\frac{\partial r}{\partial f}}\right)^{\frac{2-\gamma}{\gamma-1}}\left(\frac{e''(f)}{A_GNS\frac{\partial r}{\partial f}} - \frac{\partial^2 r}{\partial^2 f}\times\frac{A_G\gamma NS\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}\frac{\partial r}{\partial f}}{\left(A_GNS\frac{\partial r}{\partial f}\right)^2}\right) - NS\frac{\partial r}{\partial f} < 0$$

Since the inverse function of decreasing function is also decreasing, this paper proves that:

$$\frac{\partial f^*}{\partial B} < 0$$

To prove $\frac{\partial L}{\partial B} > 0$

Equation (6) determines the equilibrium distribution of population L across cities:

$$\bar{U} = \frac{\psi(NSf + d)}{L} + \left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma} - e(f)$$

The first order condition with respect to budget B is:

$$-\frac{\psi(NSf + d)}{L^2}\frac{\partial L}{\partial B} + \frac{\psi NS}{L}\frac{\partial f}{\partial B} + A_G\gamma NS\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}\frac{\partial r}{\partial f}\frac{\partial f}{\partial B} + A_G\gamma\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}$$
$$-\gamma\frac{A_GNSr + A_GB}{L^2}\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}\frac{\partial L}{\partial B} - e'(f)\frac{\partial f}{\partial B} = 0$$

$$\frac{\partial L}{\partial B}\times\left[\frac{\psi(NSf + d)}{L^2} + \gamma\frac{A_GNSr + A_GB}{L^2}\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}\right] = A_G\gamma\left(\frac{A_GNSr + A_GB}{L}\right)^{\gamma-1}$$

Therefore:

$$\frac{\partial L}{\partial B} > 0$$

To prove $\frac{\partial p_h}{\partial B} > 0$

Housing price is given by the following equation:

$$p_h = \frac{wL}{NSf + d} - \frac{1-\beta}{\beta}$$

The first order condition with respect to budgetary revenue is:

$$\frac{\partial p_h}{\partial B} = \frac{w}{NSf + d}\frac{\partial L}{\partial B} - \frac{NSwL}{(NSf + d)^2}\frac{\partial f}{\partial B}$$

$$\frac{\partial L}{\partial B} > 0, \frac{\partial f}{\partial B} < 0$$

Therefore:

$$\frac{\partial p_h}{\partial B} > 0$$

To prove $\frac{\partial q_h}{\partial B} < 0$

Housing consumption is given by the following equation:

$$q_h = \frac{NSf + d}{L}$$

The first order condition with respect to budgetary revenue is:

$$\frac{\partial q_h}{\partial B} = -\frac{NSf + d}{L^2}\frac{\partial L}{\partial B} + \frac{NS}{L}\frac{\partial f}{\partial B}$$

$$\frac{\partial L}{\partial B} > 0, \frac{\partial f}{\partial B} < 0$$

Therefore:

$$\frac{\partial q_h}{\partial B} < 0$$

Endogenous budgetary revenue:

If I endogenize budgetary revenue as the local income tax:

$$B = L\tau w$$

Population size under the spatial equilibrium will be determined by:

$$\overline{U} = \frac{\psi(NSf + d)}{L} + \left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma} - e(f)$$

Local government designs FAR limit f to maximize population. The first order condition is:

$$-\frac{\psi(NSf + d)}{L^2}\frac{\partial L}{\partial f} + \frac{\psi NS}{L} + A_G \gamma NS\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1}\frac{\partial r}{\partial f} - \gamma\frac{A_G NSr}{L^2}\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1}\frac{\partial L}{\partial f} - e'(f) = 0$$

At the optimal point, $\frac{\partial L}{\partial f} = 0$, therefore:

$$\frac{\psi NS}{L} + A_G \gamma NS\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1}\frac{\partial r}{\partial f} - e'(f) = 0 \quad (18)$$

$$\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1} = \frac{e'(f) - \frac{\psi NS}{L}}{A_G NS \frac{\partial r}{\partial f}}$$

$$w = \frac{1}{A_G \tau}\left(\frac{e'(f) - \frac{\psi NS}{L}}{A_G NS \frac{\partial r}{\partial f}}\right)^{\frac{1}{\gamma-1}} - \frac{NSr}{L\tau}$$

The derivative of w with respect to FAR f is thus:

$$\frac{\partial w}{\partial f} = \frac{1}{A_G \tau(\gamma-1)}\left(\frac{e'(f)-\frac{\psi NS}{L}}{A_G NS\frac{\partial r}{\partial f}}\right)^{\frac{2-\gamma}{\gamma-1}}\left(\frac{e''(f)}{A_G NS\frac{\partial r}{\partial f}} - \frac{\partial^2 r}{\partial^2 f} \times \frac{e'(f)-\frac{\psi NS}{L}}{\left(A_G NS\frac{\partial r}{\partial f}\right)^2}\right) - \frac{NS}{L\tau}\frac{\partial r}{\partial f} \quad (19)$$

$$\frac{\partial r}{\partial f} > 0, \frac{\partial^2 r}{\partial^2 f} < 0$$

$$e'(f) > 0, e''(f) > 0$$

$$\gamma - 1 < 0$$

Plug equation (18) into equation (19):

$$\frac{\partial w}{\partial f} = \frac{1}{A_G \tau(\gamma-1)}\left(\frac{A_G \gamma NS\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1}\frac{\partial r}{\partial f}}{A_G NS\frac{\partial r}{\partial f}}\right)^{\frac{2-\gamma}{\gamma-1}}\left(\frac{e''(f)}{A_G NS\frac{\partial r}{\partial f}} - \frac{\partial^2 r}{\partial^2 f} \times \frac{A_G \gamma NS\left(\frac{A_G NSr}{L} + A_G \tau w\right)^{\gamma-1}\frac{\partial r}{\partial f}}{\left(A_G NS\frac{\partial r}{\partial f}\right)^2}\right) - \frac{NS}{L\tau}\frac{\partial r}{\partial f}$$

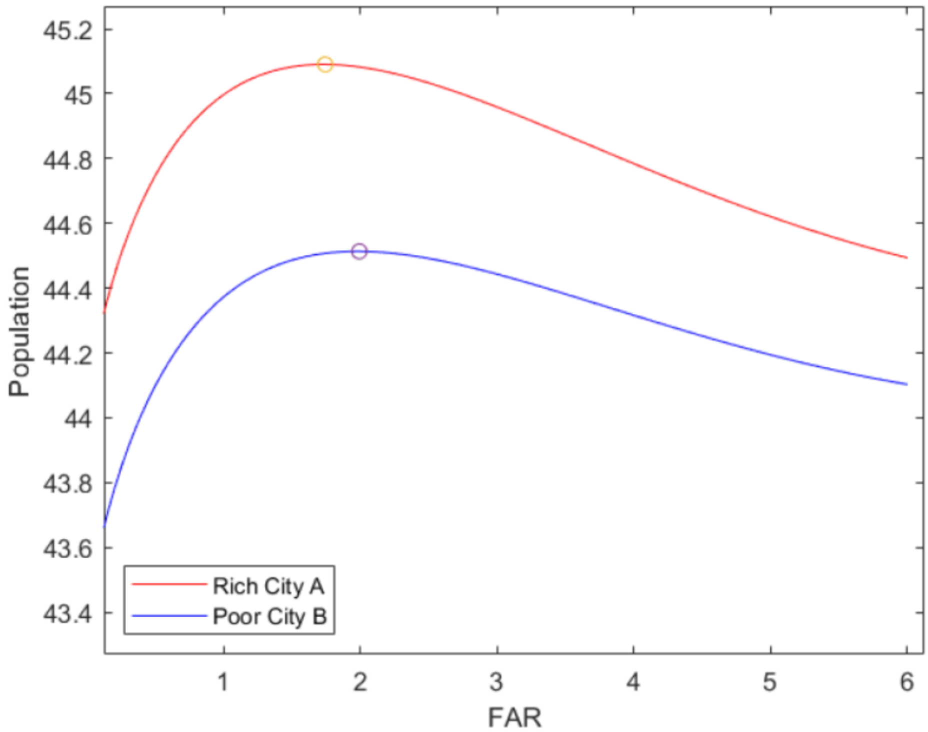
$$< 0$$

Since the inverse function of decreasing function is also decreasing, this paper proves that:

$$\frac{\partial f^*}{\partial w} < 0, \frac{\partial w}{\partial A} > 0, \frac{\partial f^*}{\partial A} < 0$$

Graphical Illustration

In this sub-section, I visualize the main proposition of the theoretical framework. The figure below shows the relationship between population and FAR upper limit. The only difference between cities A and B is that city A has more budgetary revenue. Both the red and the blue curves suggest that there exits an optimal FAR limit to maximize local population size. When FAR limit is relatively low, the positive supply and fiscal effects will exceed the negative externality effect, and there is a positive correlation between FAR limit and population. However, when FAR limit is higher than the optimal point, the negative externalities will outpace the positive effects and reduce the population size of the city. The figure below also shows that the optimal FAR limit in richer city A is lower than the optimal FAR limit in poorer city B.



Appendix D: Econometric Appendix

Bias Caused by Reverse Causality

If there is no concern of the reverse causality and omitted variables, β_1 from the following equation will be the unbiased estimate of the budgetary revenue's impact on FAR limit.

$$FAR = \alpha_1 + \beta_1 Budget$$

However, suppose that local budgetary revenue consists of $Budget_1$, which is exogenously determined, and $Budget_2$, which represents local taxes that are correlated with land value and thus FAR limits. As the proof below shows, $\frac{\beta_1}{1-\beta_1\beta_2}$ will represent the OLS estimate of the impact of budgetary revenue on FAR limit, and this coefficient will underestimate the unbiased impact. This is in line with the main empirical findings that the IV estimate is more negative compared with the OLS estimate.

$$Budget = Budget_1 + Budget_2$$

$$Budget_2 = \alpha_2 + \beta_2 FAR$$

$$FAR = \alpha_1 + \beta_1 (Budget_1 + \alpha_2 + \beta_2 FAR)$$

$$FAR = \frac{\alpha_1 + \beta_1 \alpha_2}{1 - \beta_1 \beta_2} + \frac{\beta_1}{1 - \beta_1 \beta_2} Budget_1$$

$$\beta_1 < 0, \beta_2 > 0, \ \frac{\beta_1}{1-\beta_1\beta_2} > \beta_1$$

Bias Caused by Omitted Variable

Another endogeneity concern is caused by uncontrolled local confounding factors. For instance, the density of people living in the pre-existing informal housing will increase the resettlement costs for land acquisition (Fu and Somerville, 2001), and local governments might design higher FAR limits to compensate for these residents. The literature also suggests a positive impact of informal housing on accommodating migrant inflows (Niu *et al.*, 2020), so the density of informal housing might have a positive impact on local economy and budgetary revenue.

If I can control for the population density of the pre-existing informal housing in the following specification, β_1 would be the unbiased estimate of the budgetary revenue's impact on FAR limit.

$$FAR = \alpha_1 + \beta_1 Budget + \beta_2 Density$$

However, the population density of informal housing is not controlled in the main specification due to data availability, and the biased estimate $(\beta_1 + \frac{\beta_2}{\beta_3})$ will therefore underestimate the negative effect of local budgetary revenue. This is in line with this paper's OLS and IV results.

$$Budget = \alpha_2 + \beta_3 Density$$

$$FAR = \alpha_1 + \beta_1 Budget + \frac{\beta_2}{\beta_3}(Budget - \alpha_2)$$

$$FAR = \alpha_1 - \frac{\beta_2 \alpha_2}{\beta_3} + (\beta_1 + \frac{\beta_2}{\beta_3})Budget_1$$

$$\beta_1 < 0, \beta_2 > 0, \beta_3 > 0, \beta_1 + \frac{\beta_2}{\beta_3} > \beta_1$$

Chapter 3

Why Delay?

Understanding the Construction Lag, AKA the Build Out Rate

(with Michael Ball, Paul Cheshire, and Christian Hilber)

1. Introduction

Theoretical models of residential and commercial real estate markets usually incorporate ‘construction lags’ – the time it takes from the start to the completion of a construction site. Construction lags are crucial to explain property price dynamics for example. Despite the importance of construction lags for real estate markets and property price dynamics, however, the term is typically used in the abstract. It is essentially a ‘black box’ and we know very little about the determinants.

In this paper, we employ a unique dataset of 110,000 residential sites consisting of one or more buildings, themselves representing one or multiple homes (dwellings), in England between 1996 and 2015 to explore what determines the rate of construction of these residential sites and of the individual homes on them after they are initially granted planning permission. Our sample covers over 70% of all residential developments for that period and records detailed site-level information. We first develop a theoretical framework to study the process of developers’ decision-making about how fast to build out sites. Our model predicts that positive local demand shocks increase the site build-out rate, because as demand rises developers expect a more rapid flow of sales with less downward pressure on prices. However, the impact is predicted to be weaker in places with more restrictive planning, in land-constrained already built-up areas, and in localities with a high market concentration of developers. We then empirically test for the theory’s predictions, combining our individual level site data with data for local authorities and with other controls.

The estimation employs an instrumental variable- and multiple fixed effects-strategy. The main empirical challenge is the endogeneity of both local demand shocks and housing supply constraints, as they are highly correlated with the state of local housing markets. To overcome this concern, we first employ a multiple-fixed-effect strategy and control for time-invariant features of local planning authorities and developers, and account for time trends and the seasonality of real estate development. A Bartik-type predicted local employment change measure (annualised) is used as a proxy for local demand shock, because this measure is likely to be orthogonal to local housing market conditions. Following Hilber and Mense (2021), four instrumental variables are used to address the endogeneity concerns of housing supply constraints.

In line with the theory's prediction, our empirical estimation indicates that positive local demand shocks speed up the build-out rate. Our preferred specification suggests that a 10-percentage point increase in housing demand increases the speed of construction by 36%, and the speed of construction decreases by 13%, 25%, and 23% respectively, if regulatory, geographical constraints, or market concentration increase by one standard deviation, all else equal. These effects are thus not only statistically significant but also quantitatively very meaningful. We then conduct a counterfactual analysis and find that if we reduced both supply-side constraints and the developer market power by one standard deviation, the construction duration in the 'average' English local planning authority in 2015 would have been 19 percent shorter. We carry out several robustness checks and the results are consistent with the baseline findings. We also conduct additional estimates exploring the impacts of mixed private and social housing, site size (total number of dwellings), type of building structure, and 'planning gain' Section 106 agreements on site build-out rates.

Our findings are important for the following reasons: first, to the best of our knowledge, this is the first study to explore the determinants of the 'construction lag' by applying a rigorous identification strategy and estimating on a large and representative dataset for individual dwellings and construction sites. Our paper thus differs from previous research based on interviews and surveys and provides unique insights into this research topic.

Second, our empirical findings indicate that both a more restrictive local application of the planning system and local market concentration delay the rate of site build-out. Better understanding the role of these factors, gives new insight into the determinants of the construction lag, a crucial concept for explaining real estate cycles and price dynamics.

Third, this paper contributes yet more evidence on the complex and unexpected effects the very restrictive British planning system has on the supply side of the housing market and developer behaviour. Concerns have been expressed that a significant element in the slow rate of conversion of planning permissions into completed houses is a failure of the house building industry (Office of Fair Trading 2008 & Letwin 2018) and this contributes to reduced housing supply. Our results suggest that far from there being a monopolistic conspiracy to leave planning permissions unbuilt, tighter restrictiveness itself directly contributes to slower buildout of those permissions.

We are not the first to study the determinants of build-out rates. Most of the literature on the housebuilding industry is atheoretical and the most common methods used in previous studies are semi-structured interviews and survey questionnaires (see Payne *et al.* 2019 for a summary) using samples of unknowable representativeness. For instance, Assaf and Al-Hejji (2006) use a field survey including contractors, consultants, and owners in Saudi Arabia to study the causes of construction delay. They find different potential causes such as changes in design or specification during construction, delay in progress payments, ineffective planning and scheduling by contractors, poor site management and supervision by contractors, and so on. Chan and Kumaraswamy (1997) conduct a survey to evaluate the relative importance of 83 potential delay factors for construction projects in Hong Kong. They find five principal factors: poor risk management and supervision, unforeseen site conditions, slow decision making, client-initiated variations, and work variations. One exception in terms of research methodology in the literature is Dursun and Stoy (2012), who use 1,695 observations in Germany to study the determinants of construction duration. They find that gross external floor area and cost of construction works are the major variables. Their findings also indicate that type of facility, project location, availability of construction area, and market conditions also have significant impacts on construction duration. The literature to date contains few attempts to apply a comprehensive dataset and a rigorous econometric method to explore the determinants of construction lag conceiving of that as a rational response to specific economic and regulatory determinants as well as idiosyncratic, mainly physical, site factors.

Our study also relates to several other strands of the literature. It ties into the literature on real estate cycles and dynamics (Wheaton 1999, Head *et al.* 2014) by exploring the determinants of construction lags, a crucial concept in the theoretical models of real estate markets. It also contributes to the literature on the economic and welfare impacts of urban planning and land use regulations in the United Kingdom and the United States (Cheshire and Sheppard 2002, Turner *et al.* 2014, Hilber and Vermeulen 2016). We also address the issue of market power, estimating its impact on construction duration. This is related to the literature discussing the behaviour, costs, and consequences of monopolistic companies (for instance, Posner 1975 and Prager 1990).

Our contributions to the literature are three-fold: first, to the best of our knowledge, we are the first to employ a dataset covering the great majority of all residential developments within a country to study the determinants of the site build-out rate. Previous studies have mainly used surveys and questionnaires based on tens to hundreds of observations. Our dataset allows us to employ a rigorous instrumental- and multiple fixed effects- strategy to identify the determinants of construction lag so we can be sure the results are representative and statistically robust. Second, the unique features of the British planning system and the housebuilding industry provide us with an ideal institutional setting to study the impacts of regulatory constraints and market concentration on construction duration. Our work thus provides additional insight into the causes of construction delay because previous studies mainly discuss factors such as construction costs, site management or weather. Third, we conceive of the rate of build out as something determined not just by physical or administrative factors but as subject to a firm's choice and profit maximisation. So, we propose a simple theoretical model and test for the theory's predictions. This contributes to the literature on the theoretical explanation of construction lags.

The rest of this paper is structured as follows. Section 2 describes the details of the British planning system and policy concerns about 'delays' and proposes a simple theory to guide our empirical analysis. Section 3 presents the data sources and descriptive statistics. Section 4 discusses our empirical strategy in detail and presents our main results for the determinants of construction lags. Section 5 concludes.

2. Background and Theory

2.1. The British planning system, supply constraints and policy concerns about 'delays'

Many factors explain the rate at which residential construction sites get built-out. Even if only one house is planned for and is personally commissioned, still the start may be delayed, or construction paused. This could be because of weather or difficulties with the supply of materials or labour, financing or just idiosyncratic factors. If the development is a commercial scheme of several houses then even more factors may influence the speed of construction, both of individual houses and of the 'build-out rate' of the site as a whole. The aim of this paper is to investigate the factors in England which systematically influence the rate of construction on any given site. One of these, and, as we will show, an important one, is the unusual nature of the British land use planning system.

The fundamental framework of the British planning system is still as set in the 1947 Town and Country Planning Act. This redefined the legal concept of freehold property rights, transferring – expropriating – the right to develop land or property from the owner to the state. How the system was implemented in practice was that any intended development required permission to be granted by the Local Planning Authority (LPA). For any legally defined development proposal, the would-be developer has to apply for permission to the LPA and for bigger developments this will often be in two stages: an application for 'outline' permission to establish the principle and then, if that succeeds, applications in detail (so-called 'reserved matters'). Building cannot legally commence until all aspects of a development have been approved and Local Plans (when they exist – see below) provide only schematic outlines of requirements enabling much leeway for judgement, which generates significant costs and uncertainty over outcomes for applicants.

Developments may require a series of permissions as they are built out (Ball, 2010). For example, any proposed changes to a project subsequent to initial approval may require re-submitting the full proposal in its revised form. In addition, for the past thirty years, S106 Agreements[52] may be negotiated to provide a payment in kind for the granting of planning permission. The terms of these 'planning obligations' may be revisited as new home sales proceed. Large schemes built over a number of years will typically be built in phases with reserved matters only settled prior to the start of each

[52] These were introduced in Section 106 of the 1990 Town and Country Planning Act, hence Section 106 Agreements. Such agreements are individually negotiated with developers at the time of their application and are now most commonly used to make planning permission contingent on a proportion of below market price, 'affordable' units in any development.

new phase. For these, and other reasons, there is more than one negotiation over planning permission for most major housebuilding projects and potentially even for minor ones (Ball, 2011). This leads throughout site construction to on-going interlinkages between planning and building, rather than there being an end to the planning process once permission is initially granted.

For the purposes of the British planning system ‘development’ has a legal meaning. It is not just building on green field sites, or on previously built land. It also encompasses any change of legally defined use. Categories of use - such as residential, industrial or retail - are defined in a complex set of ‘use classes’. Over time some more minor changes have become exempt from requiring permission. These so called ‘permitted development rights’ (PDR) allow some development such as converting offices or shops, in defined locations, into residential use (see Cheshire and Kaimakamis 2021). Unlike some other systems, such as the standard US Zoning system, British planning does not typically prescribe the type of housing that is developed in a locality – for example, single family only – although there may be controls – either LPA-wide or on a case-by-case basis – on heights, finishes and densities.

For the purposes of the present analysis the key characteristic of the British planning system (one shared by that of some former British territories such as New Zealand or Canada) is that it is not ‘rule-governed’ but discretionary, with the decision-making being essentially political. While LPAs are required to have an approved and up-to-date plan, less than half do; and even if there is a valid plan, decisions often do not follow it (Barker, 2006). The decision-making body for an LPA is the Planning Committee composed of local politicians. Such committees are sensitive to local feelings, especially those of voters in the wards members of the committee may represent, and are subject to fierce lobbying from local residents.

There are powerful incentives for local residents/voters to oppose development (see, for example, Hilber and Robert-Nicoud 2013, Cheshire and Hilber 2008, or Cheshire 2018) so many proposals consistent with local plans may still be rejected. Rejections are subject to a quasi-judicial appeal process involving the Planning Inspectorate and even when that process has been exhausted, a further stage of appeal is possible to the government minister responsible for the planning system. The larger a proposed development is, the more likely it is to go to appeal and the more expensive the whole process is likely to be.

This process of decision-making with respect to development contrasts sharply with 'rules-based' systems of planning typical in Continental Europe and the USA. Most countries' decisions about development are not subject to uncertainty or negotiation. In France, there is a local plan approved democratically by the local community – most often the Commune – and the role of local government is no more than to ensure that any development conforms to the local plan and design rules. That is, in most countries, decision making on development almost eliminates uncertainty as to the outcome of the decision much as in the British system of Building Regulations where evaluation simply ensures that development conforms to the rules in force.

Combined with national policies aimed at restricting land for urban development (notably Green Belt policy preserving very large zones around all major cities in which development is not allowed[53]), the empowerment of those who lose from development, including the great majority of local taxpayers, means that in economic terms there is a shortage of development pushing up the price of land for housing and housing itself. Hilber and Vermeulen (2016) developed a measure of 'planning restrictiveness', based on the proportion of applications for developments of 10 or more houses an LPA had historically rejected. They were then able to show how this restrictiveness measure translated into higher local house prices.

The overall result is that housing has become ever more unaffordable (see Cheshire, 2014) through an absolute shortage of land and a lack of timely responsiveness to changes in demand. We explore how the - by international norms - unusual, but measurable features of Britain's planning system influence the rate at which developers build out their sites in section 4 of this paper.

A further influence of the British planning system is on the firm structure of the housebuilding industry. Both the substantial fixed costs imposed on developers to operate the complexities of the system and the uncertainty and extra risk imposed on the development process have encouraged a concentration of housebuilding into large-

[53] The administrative area of London's government – the Greater London Authority area- is 159,624 ha: its Green Belt covers some 514,000 ha. Just over 22 percent of the GLA area is in the Green Belt but the great majority of London's Green Belt – some 94 percent - is outside the GLA in the South East and East of England (see Cheshire *et al.* 2014 or Cheshire and Buyuklieva 2019)

scale firms. This contrasts with the much greater dispersion of firm sizes in most other countries and their regions (Ball, 2003). Official enquiries have downplayed potential market distortions (Office of Fair Trading 2008 & Letwin 2018). By contrast, consumer-focused commentators have highlighted notable declines in product quality and consumer satisfaction as indicators of limited competition (e.g. Ali, 2019). While the planning system is not the only cause of the secular increase in market concentration, indirectly concentration has been promoted through the induced restrictions on the location and volume of housebuilding, the resultant higher land prices and the escalating costs and uncertainty of regulatory approval. Taken together, they have raised substantially the cost of new entry and undermined the viability of what was once a thriving small firm sector (Ball, 2013). A potential competitive indicator is the speed at which firms respond to demand increases, because in the absence of local competition housebuilders have little to fear from competitors grabbing their markets by building faster. So, we explore this as well in section 4.

2.2. Theory

In this sub-section, we develop a model to guide our empirical analysis on the determinants of the construction lag. Specifically, the model explores the process of developers making decision on delaying construction based on factors including housing demand, construction costs, housing supply constraints, and developer market power.

The model first illustrates how local demand shocks influence a developer's decision to delay construction. A positive local demand shock leads to a higher realised market price[54] for housing, which makes developers more likely to achieve their expected profits and thus decide on no delay to avoid additional financial costs. If local housing demand weakens so housing prices are expected to be lower, developers choose to delay construction and wait for the market to recover. The model also proposes that under suitable assumptions – made explicit below – the relative magnitude of the demand effect depends on local housing supply conditions and developer market power. For places with more restrictive housing supply and for sites constructed by developers with higher market shares, the effect of local demand shocks on the site build-out rate is weaker.

[54] Allowing for speed of sale.

Suppose that developer A produces one unit of housing and can observe the local demand level D. Developer A predicts housing price to be $h(D)$, an increasing function of D, when a unit is sold. C denotes the construction cost of housing and X denotes the supply constraints that make construction more costly (e.g. geographical obstacles and regulatory requirements that developers need to take into account when building out sites). ε is the prediction error between the realised market price for housing and the developer's expected housing price $h(D)$. Suppose that ε follows a standard uniform distribution, $\varepsilon \sim N(0,1)$, with cumulative density function F_e and probability density function f_e. The developer's profit P is thus given by:

$$P = h(D) + \varepsilon - CX$$

Suppose that there is an expected profit level $\bar{P}$ for developer A. If developer A could make profits greater than $\bar{P}$, they will not delay construction and build the housing unit as quickly as possible. We thus define the developer's decision to delay construction as an indicator function with the expected profit level $\bar{P}$:[55]

$$\mathbf{1}_{\bar{P}}(P) = \begin{cases} 1 \ \ if\ P < \bar{P} \\ \\ 0 \ \ if\ P \geq \bar{P} \end{cases}$$

At the break-even point when $P = \bar{P}$, ε equals to $\bar{\varepsilon}$:

$$\bar{\varepsilon} = \bar{P} - h(D) + XC$$

L represents the probability that developer A choose to delay the construction, and could be computed using ε's cumulative density function F_e:

$$L = F_e(\bar{\varepsilon}) = F_e[\bar{P} - h(D) + XC]$$

We could then prove that:

$$\frac{\partial L}{\partial D} = -h'(D)f_e[\bar{P} - h(D) + XC] = -h'(D)\frac{1}{\sqrt{2\pi}}exp(-\frac{(\bar{P} - h(D) + XC)^2}{2}) < 0$$

[55] This model only indirectly and crudely considers the opportunity cost of time faced by developers. Rather than optimizing the timing, by taking into account both the opportunity cost of time and the achievable sale price, developers in our setting start to build once they observe a profitable price. In future work, we will develop a dynamic framework that models the delay decision of developers more concisely.

This inequality shows that as local demand increases, the probability of developers delaying construction will decrease. In other words, positive local demand shocks will increase site build-out rates, because developers are more likely to meet their expected profit level and there is no need to delay construction and bear costs.

We further take into account the impacts of local supply constraints X on developer's decisions with respect to delaying construction:

$$\frac{\partial^2 L}{\partial D \partial X} = h'(D)(\bar{P} - h(D) + XC)\frac{C}{\sqrt{2\pi}} exp(-\frac{(\bar{P} - h(D) + XC)^2}{2}) > 0$$

This inequality shows that the impact of positive local demand shocks on build-out rates are weaker if the site is located in a place with restrictive housing supply conditions, because these supply constraints increase construction costs and make it more difficult for developers to meet their expected profits.[56]

Suppose that developer A's expected profit $\bar{P}$ is dependent on their market power M. If a developer doesn't have any market power (e.g. the developer is in a perfectly competitive housing market), $\bar{P}$ will be zero. If a developer has nearly monopolistic market power, $\bar{P}$ will be high. We can then prove that:

$$\frac{\partial^2 L}{\partial D \partial M} = h'(D)(\bar{P} - h(D) + XC)\frac{d\bar{P}}{dM}\frac{1}{\sqrt{2\pi}} exp(-\frac{(\bar{P} - h(D) + XC)^2}{2}) > 0$$

This inequality shows that the impact of local demand shocks on site build-out rates are weaker if the developer has a higher market power. This is because developers with more market power expect to sell housing units at higher prices, so they are more likely to delay the construction and wait for the best price.

Based on these inequalities, our main proposition can be formulated as:

Proposition – Positive local demand shocks make site build-out rates faster, as developers are more likely to achieve their expected profits in a market with stronger housing demand. This demand effect is weaker for projects located in areas with more supply constraints and for projects constructed by developers with a higher degree of market power.

[56] Another potential mechanism for supply constraints influencing site build-out rate is through the uncertainty of real estate development caused by restrictive planning system. We don't discuss this mechanism due to the static nature of our theoretical framework, but the overall impact of supply constraints is empirically estimated in section 4.

3. Data and Descriptive Statistics

Our empirical analysis employs unique geo-located data on construction sites in England, covering – for the years in our regression sample – the vast majority of all sites and including detailed information on the site build-out rate (i.e., the speed at which the site is developed) and site-, as well as dwelling-characteristics. The data includes both the geo-coordinates of the boundaries of the whole construction site and those of the individual dwellings within the construction site, including when each dwelling was started and when it was finished - see below.

Our main data source is the National House Building Council (NHBC), the leading provider of warranty insurance for new homes in the UK. The NHBC is a non-profit company independent of government and the construction industry. It records inspections for construction sites at key build stages and covers most residential developments in the UK since the 1990s. The NHBC dataset used in our analysis contains information on completed sites, comprising 3,261,880 dwellings, constructed in England between 1986 and 2020. The information includes dwelling start and completion dates, site locations, dwelling types (flats or houses), the number of bedrooms for each dwelling, a unique construction site identifier, a unique dwelling identifier, the site developer, and whether a housing association participates in the development. For the purpose of our analysis, a dwelling is labelled as a 'public unit' if it is either built by a housing association or jointly developed by a private developer and a housing association.[57]

To perform our empirical analysis, we aggregate the NHBC dwelling-level records to the construction site level. For each construction site, we compute the share of public dwellings to measure its public housing intensity. We also calculate the share of flats within a site to characterise its physical structure and count the number of dwellings within a site to measure its size. We employ the National Statistics Postcode Lookup Directory to match construction sites to coordinates and local planning authorities.[58]

[57] Usually as part of S106 agreements, a site once entirely belonged to a developer who then 'sells' part of the site to a housing association to build it out.

[58] We use the most frequently occurring postcode within a construction site to geocode it. In the dataset, approximately 5% of dwellings do not contain full postcode information, and we use their postcode district information to spatially link them to the corresponding local planning authority. These dwellings are not included in our site-level estimations but are included in our LPA-level estimations.

The NHBC dataset covers the vast majority of residential developments in England since the 1990s. Figure 1 illustrates, respectively, the number of NHBC dwellings, the number of NHBC private dwellings, and the total number of dwellings started in England as reported by the Ministry of Housing, Communities and Local Government (MHCLG). As shown in Figure 2, the share of all residential construction covered by the NHBC improved steadily during the late 1980s until around 2002. It reached 84% in 1996 and coverage remained very high, between 70% and 90%, all the way through until 2015. During this period, the NHBC covered essentially the entire country, that is, there was coverage in nearly all 353 LPAs in England. The coverage of the dataset starts to drop very significantly from 2015 for a technical reason: the data set only records construction sites that are fully completed. Thus, it increasingly misses dwellings on sites still under construction the more recent the year, creating a sample selection issue.

To deal with this issue we confine the sample for our baseline regression to the period from 1996 to 2015. We start with 1996 because this is the first year when the NHBC's digital recording covers much of the market (i.e., 84%). We drop the years from 2016 onwards to minimise the concern that our results might be affected by sample selection. We pick 2015 as our final year for three reasons. First, as Figure 2 illustrates, coverage of total housebuilding starts to drop dramatically from 2016 onwards. Second, and related, the average construction duration for large sites with more than 100 dwellings in our sample is 1573 days (around 4.3 years). This implies that most sites that started in 2015 were in fact completed by 2020, when the NHBC coverage stops. Put differently, our sample of NHBC dwellings in 2015 is likely to be very close to the final count, which will only be known in a few years from now. Third, and again related, picking 2015 as final year ensures our analysis is not affected by COVID-19, which, for a time, brought construction nearly to a stand-still.[59] Using data for the period from 1996 until 2015 ensures that our regression sample provides a comprehensive coverage of the market for new homes in England.

Before turning in more detail to the variables included in our regression sample, it is worth stressing the worldwide unique character of the NHBC dataset. We are unaware of any other dataset that provides a similarly comprehensive coverage of detailed

[59] As we report in Section 4.5, our results are essentially unchanged if we replicate our analysis for the period from 1996-2016 or if we, in fact, make use of more NHBC data from 1986 to 2018.

construction activity at site-level for a large country such as England. Apart from the excellent coverage of the construction market in England, the NHBC dataset has several important additional advantages. First, it contains detailed information on the size of construction sites, the structures built, locations and developers. This enables us to study build-out rates and their determinants. Second, the NHBC is an independent, non-profit organisation and its detailed inspections provide accurate records of the residential development process. Third, the large spatial variation in land use planning restrictiveness across the LPAs of England provides an ideal institutional setting to explore regulatory impacts on site build-out rates.

We define the 'construction duration' for each site as the time between its start and its completion-date.[60] Panel A of Figure 3 shows that the construction duration for most sites in our baseline sample is less than 2,000 days, although there are a few sites with a significantly longer construction duration, up to 7,720 days in the extreme.[61] Figure 4 presents the yearly average construction duration across projects of different sizes in England. The average construction lag is in general stable over time, with an exceptional increase between 2007 and 2010, likely to be explained by the financial crisis.

To measure local developer competitiveness and market power, we first draw a 10-km radius buffer for each construction site and then normalize the radius of each buffer based on each site's corresponding Travel To Work Area (TTWA) population density in 2001. The mean population density at the LPA level in 2001 is 1353 persons/km2. If a site is in an LPA with population density higher than 1353, we will adjust its corresponding radius based on its LPA population density and the national mean value. Otherwise, we compute the local market share within the 10-km radius buffer. We then compute the market share of each site project's developer within the adjusted buffer. The reason to adjust the radius of each buffer is because we want to define a local housing market with more comparable population size and potential buyers. In the

[60] For each site, the start date is observed when the first slab of site is completed and a NHBC inspection is triggered. Before the start date recorded by the NHBC, there may be some time for land preparation, infrastructure development, and the digging out and laying of the first batch of dwellings' foundations and services. The completion date is the NHBC inspection date of the last completed dwelling. There might still be some final works after that such as landscaping, roads, etc. Both definitions suggest that the actual construction duration for each site might be even longer than our dataset records.

[61] We dropped sites with construction duration fewer than 10 weeks in our baseline sample to mitigate potential measurement errors. However, as shown in the appendix Table A11, including such sites makes no difference to our results.

appendix, we report estimation results using local market share measure with no adjustment and our findings are robust.

Figure 5 presents an example. We first draw a 10-km-radius buffer for site A. Within this buffer, three construction sites (A, B, and C) all start in the same year. Both sites A and B are constructed by developer X, and site C is constructed by developer Y. Site A's local market share is then defined as the number of dwellings within sites A and B (both constructed by developer X) relative to the total number of dwellings within sites A, B, and C. Panel B of Figure 3 presents the histogram of the local market share based on our main estimation sample. We zoom in the tail part to show the distribution of high market shares. For the vast majority of site locations, the corresponding developer has a local market share of below 20%. There are however a few large sites, where the developer has a (near) monopolistic market share of up to 100%.

To capture local developer competitiveness at the LPA level, we compute two measures: a standard Herfindahl–Hirschman Index[62] and the top-5 developers' market shares, respectively, for each LPA. Figure 6 illustrates how the Herfindahl–Hirschman index and the top-5 and top-10 developers' market shares at the national level evolve over time. The figure reveals that the construction market in England is heavily influenced by large developers, with the top-10 developers producing more than 40% of all new homes, in each year between 1996 and 2015.[63] Figure 6 also shows that after the financial crisis in 2007, both the top developers' market share and the Herfindahl–Hirschman index increased, suggesting that big developers were more likely to survive the crisis and thus gain a higher market share post-crisis.

We spatially merge the NHBC dataset with data from other sources to get information about local housing demand and supply conditions, and to control for a wide range of geographical and weather conditions. We collect data about the refusal rate of major projects and S106 completions from the MHCLG. We compute the average refusal

[62] The Hefindahl-Hirschman index, calculated as the sum of the market share percentages of developer k within a buffer, is a common measure of market concentration. It can range from 0 to 1. The higher the index, the greater is the market concentration. A low value implies a competitive marketplace.

[63] Table A1 presents the market share of the top 10 developers in England between 1996 and 2015 based on our estimation sample. Top 10 developers in total account for over 49% of all new homes in England during our main estimation period. It is also worth noting that there have been extensive mergers and acquisitions amongst housebuilding firms during the sample period, leading to greater concentration, but that for computational purposes we assume named builders in the sample are independent whereas some may have been jointly owned at the time. Our concentration estimates thus should be regarded as minima.

rate at the LPA level between 1996 and 2015, the same period for our main empirical specification, as a proxy for local regulatory restrictiveness. As shown in Appendix Figure B1, LPAs in London and the Southeast region tend to have higher refusal rates for major applications and are thus more likely to have restrictive planning environments. We also use the number of S106 completions between 1996 and 2015 relative to the housing stock in 2015 at the LPA level to measure local S106 intensity. We document the spatial variation of the S106 intensity in Appendix Figure B2. We use a Bartik-type shift-share measure (i.e., the predicted local employment based on an LPA's industry composition and the national-level employment growth rates of these industries), taken from Hilber and Mense (2021), to capture shifts in local housing demand. Appendix Figure B3 presents the percentage change in predicted employment at the LPA level between 1996 and 2015. As the figure illustrates, housing demand and economic prosperity are highest in the Southeast and in Greater London. We discuss the predicted local employment measure in more details in Section 4.1.

We also use data from Hilber and Mense (2021) to measure local geographical constraints (the share of developed land relative to developable land in 1990) and to construct instruments for our supply constraint variables. We discuss the identification strategies in section 4.1. Following Gibbons *et al.* (2019), we consider the centroid of each TTWA as a proxy for the city centre and compute each geocoded site's distance to the so defined city centre. Finally, we obtain weather and soil condition data from the Met Office and the British Geology Survey respectively, and spatially merge each site with its corresponding weather and soil conditions so that we can control for within-LPA features. We adjust all the LPA-level variables to the 2001 Census LPA administrative boundaries to stay consistent in our empirical estimation.

In Panel A of Table 1, we present the summary statistics for the sample of construction sites between 1996 and 2015 in England. There are 119,003 completed construction sites in this sample. The average construction duration of build out is 594 days, and on average, there are 18 units per construction site. The average number of bedrooms per dwelling is 3, and the average share of public units within a site is 6%. Panels B and C show the summary statistics for the LPA-level variables. The average refusal rate for major projects is 19% and the average Herfindahl–Hirschman Index is 0.09. Panel D presents the summary statistics for four instrumental variables that we apply to generate exogenous variations in supply-side constraints. We discuss these variables and the details of our identification strategy in Section 4.1.

Panel C of Figure 3 presents a histogram of the number of dwellings within each site based on our baseline sample. Most sites have between 1 and 100 dwellings, though a few large sites have more than 1000. Figure 7 presents the time-series of both the average number of dwellings within each site and the number of medium and large sites relative to all sites[64] in England between 1996 and 2015. The average size of a construction site has been relatively stable between 15 and 20 units between 1996 and 2009. It increased markedly after 2009 to around 25 units. This phenomenon might be driven by the fact that big developers were more likely to survive the financial crisis.

The construction duration (Appendix Figures B4 to B5) and local market power (Appendix Figure B6) also vary substantially spatially in our regression sample. Figure B4 illustrates that it takes significantly longer for developers to build sites in the Northeast compared to the Southeast. We have two explanations for this. First, housing demand growth is much higher in the Southeast (see Appendix Figure B3). Second, consistent with developable land being scarcer and planning restrictions being tighter, more small sites tend to be built in the Southeast. Conversely, Appendix Figure B7 documents the average construction duration at *dwelling-level*, suggesting that the construction of individual units takes longer in the Southeast and Greater London than in the Northeast. While demand pressures are higher in the Southeast and in Greater London, these regions are also characterised by tight planning controls and a high degree of physical development, both potentially slowing down the construction speed per dwelling. When we explore variation within cities (Appendix Figure B5), we find that sites close to the Central Business District (CBD) tend to have shorter construction durations. Again, this might be due to stronger housing demand pressures near or in the CBD.[65] Appendix Figure B6 documents the spatial variation in developers' market power, employing a Herfindahl-Hirschman Index[66] at LPA-level. The figure reveals that the largest developers in the North of the country tend to have more market power.

[64] Medium refers to projects with 25 to 100 unit. Large refers to projects with more than 100 units (Ball, 2011).

[65] Figure B6 presents the spatial variation of construction duration within big cities such as London, Manchester, Birmingham using the reverse distance smoothing technique.

[66] We compute the Herfindahl–Hirschman Index by squaring the market share of each developer within each LPA and then summing the resulting numbers for each LPA between 1996 and 2015. For example,

Appendix Figure B8 finally shows significant seasonality for the start and completion months of construction sites. Both Panels A and B suggest that more projects start and complete during the summer (especially in June), while comparably few projects start and complete in December and January. We control for the seasonality of residential development by including both the start month and the completion month fixed effects in our baseline specification.

4. Empirical Analysis

4.1. Endogeneity Concerns and Identification Strategy

The focus of our empirical analysis is to test our theoretical proposition that an increase in housing demand, all else equal, will speed up construction, but less so in locations where supply constraints are more restrictive and competition in the development sector is lower. To test these hypotheses, we interact a variable that captures housing demand with supply constraint-variables and our measure of developer competitiveness, respectively.

As previously noted, our local housing demand shifter is the annualised change of a Bartik-style shift-share measure – the predicted local employment – taken from Hilber and Mense (2021). The shift-share measure is derived by transforming time-series industry variation ('shift') at the national level into local shocks based on local industry composition ('share') at the LPA level in 1971, pre-dating our main sample period by 25 years. The predicted employment arguably introduces an exogenous demand shock to local housing markets, as both the 'shift' and the 'share' variables are likely to be orthogonal to the state of the local housing market between 1996 and 2015. We then compute the annualized change in local predicted employment (ACLE) for each site as:

$$ACLE_{jt^st^e} = Annualised\ Change\ in\ Local\ Employment_{jt^st^e} = \frac{predicted\ employment_{jt^e} - predicted\ employment_{j(t^s-1)}}{(t^e - t^s + 1) \times predicted\ employment_{j(t^s-1)}}, \quad (1)$$

whereas t^s and t^e refer to the start- and completion-year of each construction site. $ACLE_{jt^st^e}$ represents the annualised change in the predicted local employment in local

for an LPA with 3 developers with market shares of 20, 20, and 60 percent respectively, the HHI is 0.44 ($0.2^2 + 0.2^2 + 0.6^2 = 0.44$).

authority j between year t^s and year t^e. This variable thus captures the annualised demand shock for each construction site during its development period. Since the start- and completion-years vary from site to site, our demand measure may vary within LPA across sites.

To capture developer competitiveness, we use both the local market share measures and the Herfindahl–Hirschman index, discussed in Section 3. One concern related to our measure of developer competitiveness is that tight land use regulations may themselves impose a hurdle that requires a complex process and corresponding skills to negotiate planning permission. This in turn may prevent small developers from entering the market. To address this endogeneity concern, all our specifications include LPA fixed effects. These control for all time-invariant characteristics at the local level including regulatory restrictiveness so our estimate of the impact of developer competitiveness is conditional on the local regulatory restrictiveness.

Our measure of regulatory restrictiveness is the average refusal rate of major residential planning applications between 1996 and 2015 derived from the MHCLG. The refusal rate of 'major applications' (i.e., applications of projects consisting of ten or more dwellings) is the standard measure used in the literature to capture regulatory restrictiveness in Britain – see Cheshire and Sheppard (1989), Bramley (1998), or Hilber and Vermeulen (2016). We compute the average refusal rate between 1996 and 2015 to mitigate the concern regarding the pro-cyclical nature of local planning decisions. The other supply constraint measure, the share of developable land already developed in 1990, is taken from Hilber and Vermeulen (2016) and is used as a proxy for local physical obstacles.

Although our housing demand shifter, ACLE, is likely to be exogenous to local housing market conditions, two of our supply constraints variables, the refusal rate and the share developed measures, are arguably endogenous. The concern is that these measures are correlated with local housing demand (see e.g. Davidoff 2016), which has a direct impact on the site build-out rate. Moreover, as local planning decisions are the outcome of a political economical process and are shaped by homeowners, developers, and politicians (see e.g. Hilber and Robert-Nicoud 2013), local planning refusal rate is likely to be correlated with developer characteristics. For example, larger developers may be better equipped to deal with restrictive planning authorities and may be more likely to gain planning approval. These developers may also have specific construction and project management techniques that could affect the site build-out rate. These larger developers can thus 'prosper' in more restrictive LPAs and crowd out smaller ones. We first mitigate the endogeneity concern by controlling for developer fixed effects, effectively comparing within-developer variation in construction duration across sites. However, confounding factors such as time-varying developer characteristics or local political features might still bias our estimates. If more capable developers were able to reduce both the likelihood of rejected planning applications and the site construction duration, the OLS estimate of the impact of planning restrictiveness would be biased. In addition, the 'share developed' variable is potentially endogenous as it is determined by contemporaneous demand and supply conditions.

To address these endogeneity concerns, we follow Hilber and Vermeulen (2016) and Hilber and Mense (2021) and employ an instrumental variable strategy. We employ three instrumental variables for the refusal rate. Our first instrument is the LPA share of greenbelt land in 1973, 23 years prior to the start of our sample period. Greenbelts in England represent major obstacles to new development. We would expect that those LPAs that were assigned a large share of greenbelt land back in 1973 were also those LPAs with strong cohorts of Not-in-My-Backyard (NIMBY)-residents who benefited from nice views and protected asset values. These LPAs are likely to have more restrictive planning generally, so the share greenbelt land can be expected to be positively correlated with our refusal rate measure. Yet, the historic share of green belt land should not affect contemporaneous changes in the speed of the build-out rate other than through regulatory restrictiveness.

Our instruments two and three for the refusal rate were initially proposed by Hilber and Vermeulen (2016). The second instrument is derived from a reform of the English planning system in 2002, which imposed a speed-of-decision target for major developments onto LPAs but did not alter an LPA's ability to refuse planning applications. LPAs therefore had the option to substitute one form of 'penalised' restrictiveness (not meeting a delay target) with another 'non-penalised' form (refusing planning applications). Hilber and Vermeulen (2016) show that changes in the delay rate and changes in the refusal rate were uncorrelated before it became clear that the delay rate targets are introduced, the two measures however become strongly negatively correlated afterwards. Our identifying assumption is that the reform had a differential impact on less and more restrictive LPAs: the most restrictive LPAs should have had the strongest incentives pre-reform (measured between 1994 and 1996) to delay residential applications and the strongest incentives post-reform (measured between 2004 and 2006) to reduce their delay rate by refusing more of them. We would not expect the change in the delay rate pre- vs. post-reform to directly (other than through regulatory restrictiveness) explain changes in contemporaneous build out-rates.

The third instrument is the vote share of the Labour party in the 1983 General Election (derived from the British Election Studies Information System). On average, voters of the Labour party have below-average incomes and housing wealth and they are significantly more likely to rent. Hence, we expect this group to care less about the protection of housing wealth and to be more likely to vote for politicians who favour a laxer planning environment. This suggests a negative correlation between the Labour vote share and local planning restrictiveness. Our identifying assumption is that the share of Labour votes affects construction duration only through its impact on local restrictiveness, after controlling for LPA-, developer-, and time-fixed effects.

The share of developable land developed in 1990 is potentially endogenous to local demand conditions. We adopt the strategy proposed by Hilber and Vermeulen (2016) and instrument the share of developed land in 1990 with the historic population density in 1911. The rationale is that population density in 1911 can be expected to be strongly correlated with time-invariant local amenities and the inherent productivity of a place, which in turn can be expected to be positively correlated with the share of developed land in 1990. Meanwhile, the direct impacts of these amenities and productivity on construction duration will be captured by the LPA-fixed effects. Historic population density can therefore be expected to only influence the site build-out rate through affecting the scarcity of developable land in 1990.

4.2. Econometric Specifications

Our baseline specification at the construction site level is:

$$Log(construction\ duration)_{ijkt^st^em^sm^e} = \beta_1 ACLE_{jt^st^e} + \beta_2 ACLE_{jt^st^e} \times \overline{refusal\ rate_j} + \beta_3 ACLE_{jt^st^e} \times \%developed_j + \beta_4 ACLE_{jt^st^e} \times market\ share_{ik} + X_i + Z_j + \rho_k + \delta_{t^s} + Y_{m^s} + \omega_{m^e} + \varepsilon_{ijkt^st^em^sm^e} \quad (2)$$

$Log(construction\ duration)_{ijkt^st^em^sm^e}$ represents the construction duration for site i, which is developed by developer k in local planning authority j, starts in month m^s of year t^s, and completes in month m^e of year t^e. $ACLE_{jt^st^e}$ measures the annualized change in local employment in local authority j between year t^s and year t^e. We interact $ACLE_{jt^st^e}$ with the average refusal rate of major planning applications in LPA j, $\overline{refusal\ rate_j}$, the share of developable land already developed in LPA j in 1990, and developer k's local market share $market\ share_{ik}$.

All three measures of interest are in standardized form (i.e., normalized to the mean being equal to zero and the standard deviation being equal to one), so that the interpretation of the coefficients $\beta_1,\dots,\beta_4$ is straightforward: β_1 captures the impact of a labour demand shock on the site build-out rate in an LPA with average supply constraints and developer competitiveness. The coefficients $\beta_2,\dots,\beta_4$ capture the change in the impact of the local demand shock when the housing supply constraint or developer monopoly power increases by one standard deviation. We instrument for the interaction of the refusal rate and the interaction of the share developed by the interactions of the annualised change in local employment with the four instrumental variables discussed above.

In addition, we control for a wide range of site-level characteristics X_i including number of dwellings, share of public dwellings, share of flats, distance to CBD, average number of bedrooms per unit, and a dummy denoting whether there is a change of developer.[67] We include LPA fixed effects Z_j, developer fixed effects ρ_k, year fixed effects δ_{t^s}, start month fixed effects m^s, and completion month fixed effects m^e to control for time-invariant features at LPA-level, time-invariant characteristics for each developer (e.g. project management ability and the speed of decision making), the national macro trend, and the seasonality of real estate development respectively. Finally, we include weather and soil conditions[68] at the site level to control for within-LPA differences in geological and weather conditions. We cluster our standard errors at LPA-level to account for potential spatial correlation in construction duration.

[67] Occasionally we observe the change of developer within a construction site. This can be driven by either mergers and acquisitions of developers or the split of a large construction site into several smaller projects.

[68] We control for soil texture fixed effects which include 38 different types such as clayey, loam, peat, sand, and so on.

Our key explanatory variables of interest are all LPA-specific and one could argue that because we exploit variation in these variables only at LPA-level, there is little benefit to estimating our baseline specification at site-level. Besides, LPA is an important geographical unit for the local planning system and the estimation results at the LPA level are also quantitatively meaningful. Thus, to test for the robustness of our site-level estimation, we also estimate the impacts of supply constraints and developer monopoly power on the construction duration-local labour demand shock (LLDS) elasticity at the LPA level:

$$Log(construction\ duration)_{jt} = \beta_1 LLDS_{jt} + \beta_2 LLDS_{jt} \times \overline{refusal\ rate_j} + \beta_3 LLDS_{jt} \times \%developed_j + \beta_4 LLDS_{jt} \times Herfindahl_j + X_{jt} + \delta_t + Z_j + \varepsilon_{jt} \quad (3)$$

$LLDS_{jt}$ is the natural logarithm of predicted local employment in local authority j in year t. $Herfindahl_j$ denotes the Herfindahl–Hirschman index in local authority j. We aggregate construction sites starting in year t in LPA j to create time-varying variables at the LPA level. $Log(construction\ duration)_{jt}$ is the average of construction duration in local authority j in year t, which is computed based on construction sites starting in year t in LPA j. X_{jt} denotes time-varying LPA-level features including the average size of sites, the average share of public dwellings on a site, the average share of flats, and the average number of bedrooms in local authority j in year t. We include LPA fixed effects and year fixed effects in the specification to control for time-invariant LPA-level unobserved features and the nation-wide macro trends respectively. We instrument for the two supply constraint interaction variables following the strategy discussed above. We also conduct a counterfactual analysis based on the estimates from specification (3) and the details will be discussed in section 4.4.

4.3. Baseline Estimation Results

Key explanatory variables (supply constraints and developer market power)

Table 2 summarizes our main findings for estimating equation (2). Columns (1) to (3) report results for naïve OLS specifications, cumulatively adding controls, first only LPA-, year-, and month- fixed effects plus site characteristics, then additionally developer fixed effects, and, finally, micro-location weather and soil condition controls. All observations are clustered at the 2001 local planning authority level to account for potential spatial autocorrelation in construction duration.

The coefficients on the impacts of ACLE on construction duration are highly statistically significant and negative in all three columns. Column (3) implies that in an LPA with average supply constraints and average developer competitiveness, a 10-percentage point increase in local housing demand decreases construction duration by 36%. The coefficients on the ACLE-interactions with the refusal rate and the share developed-measure are positive and statistically significant in most specifications. Column (3), the most rigorous of the OLS specifications, implies if an 'average LPA' observes a one standard deviation increase in its refusal rate, a 10-percentage point increase in local housing demand will decrease the construction duration only by 22% instead of 36%. The coefficient on the ACLE-local market share interaction is also statistically significant and positive in all three specifications. The coefficient in column (3) is 1.16, implying that an LPA with a one standard deviation higher market concentration than the average LPA (all else equal) will see the speed of construction decrease by only 25% instead of 36% as a consequence of a 10-percentage point increase in local housing demand.

The OLS specifications ignore endogeneity concerns related to the local regulatory restrictiveness and the share developed land measures. In columns (4) to (6), we estimate the same regression by Two-Stage Least Squares (2SLS), instrumenting the refusal rate-ACLE and share developed-ACLE interactions. Similar to the OLS specifications reported in columns (1) to (3), we sequentially add developer fixed effects and micro-location weather and soil conditions. Consistent with the OLS estimates, the coefficients on the refusal rate-ACLE, the share developer-ACLE, and the local market concentration-ACLE interactions are all positive and highly significant. Moreover, the Kleibergen-Paap F statistics do not indicate that weak identification is a problem.

Our most rigorous specification reported in column (6) suggests that a 10-percentage point increase in housing demand increases the speed of construction by 36%. The estimated coefficient is almost identical to the one reported in column (3). Moreover, the estimated coefficients in column (6) reveal that the speed of construction only decreases by 13% (instead of 36%), 25%, and 23% respectively, if the refusal rate, the share developed and market concentration increase by one standard deviation, all else equal. These effects are thus not only statistically significant but also quantitatively very meaningful. We explore the quantitative effects further, in Section 4.4 below. The fact that the estimated coefficient, especially of the refusal rate measure, is larger in magnitude in the IV- than the OLS-specifications is moreover consistent with our argument that the confounding factors in the OLS specification are likely to lead to biased estimates of the impact of planning restrictiveness. We also include a wide range of site-level controls in column (6) and we will discuss the estimated results for these controls in detail in below.

We report the first-stage regression results, corresponding to columns (4) to (6) of Table 2, in Table 3. In all first-stage regressions, the share of greenbelt land in 1973, the reform-based change in the delay rate, and the Labour party vote share correlate strongly and in expected ways with the refusal rate of major planning applications. In addition, the historic population density in 1911 has a positive and statistically significant correlation with the share of developable land already developed in 1990.

Table 4 presents our OLS and IV estimates at the local planning authority level. The estimates for our variables of interest are consistent with the construction-site level findings. Table 4 also documents that it takes longer to build sites with more dwellings, more bedrooms, and more flats, and it takes less time to build sites with more public units, presumably as this facilitates interactions with LPAs. The corresponding first-stage results at the local authority level are as expected and are shown in Table 5. Kleibergen-Papp F-statistics again do not reveal a problem with weak identification.

Additional controls

To provide further insights into the determinants of construction lags, we report the coefficients of our additional control variables in Appendix Table A2. Columns (1) to (3) report the coefficients of our OLS estimates and columns (4) to (6) report the results when we instrument for the potentially endogenous supply constraints. Our estimates for the control variables are robust across all six specifications. Focusing first on site characteristics, we find that both the number of dwellings per site and the number of bedrooms per unit significantly increase construction duration. The estimated coefficient for the share of flats is negative and statistically significant, suggesting that conditional on the number of dwellings, it takes longer to build single-family houses rather than flats. We also find, not surprisingly, that if a site experiences a change in developer during the construction period, its build-out rate will be slower. This is likely due to potential planning adjustments and rearrangements and interruptions caused by the replacement of the developer. The estimated coefficient for the share of public units is negative and statistically significant, suggesting that sites with more public units tend to be constructed faster. This finding is plausible as public housing is typically constructed by housing associations, which might be able to more easily cut through 'red tape'. Besides, housing associations have guaranteed funding, use contractors with time penalties to build, and usually have 'pre-let' tenants to fill completed dwellings and so don't have to wait for market conditions in order to sell. We also find that spatial controls and micro-location weather characteristics matter in expected ways. Table A2 shows that there is a positive and statistically significant correlation between a site's distance to CBD and the construction duration (in line with Appendix Figure B5), as housing demand tends to be higher for places close to city centre. Meanwhile, we observe a negative and statistically significant association between temperature and construction duration, and a positive and statistically significant association between wind speed and construction duration. Overall, these latter findings suggest that spatial variation in weather conditions matters even within relatively small geographical units. That is, our analysis controls for LPA fixed effects and hence, our weather variables only exploit variation within LPAs.

4.4. Quantitative Analysis

We base our counterfactual analysis on the TSLS specification reported in Table 4. Our preferred specification is the most rigorous one reported in column (4). The specification yields a prediction of construction duration conditional on the local labour demand shock, supply constraints, developer market power as measured by the Herfindahl-Hirschman index, as well as LPA and year fixed effects.

We first obtain counterfactual scenarios by predicting local construction duration with supply constraints and developer market power set to zero sequentially. We then also remove the independent effect of the LLDS, in order to identify the counterfactual construction duration holding constant all relevant local demand and supply measures. This exercise allows us to understand the quantitative importance of our variables of interest.

Removing all supply constraints and creating a setting with perfect competition in the residential construction market are of course unrealistic scenarios in practice. Hence, we explore an alternative exercise, where we remove one standard deviation of each of the two supply constraints measures and of the market power measure, sequentially. We first conduct this exercise for each LPA separately and then take the average of the predicted construction duration over all local authorities to derive a counterfactual scenario for the 'average' English LPA. To explore the relative importance of our variables of interest, we also conduct two exercises by separately removing supply constraints and market power, and by separately lowering these variables by one standard deviation.

The results of these quantitative exercises are summarised in Table 6. The corresponding Figures 8 and 9 illustrate the predicted construction duration between 1996 and 2015 for the 'average' English LPA under two scenarios: variables of interest set to zero and reduced by one standard deviation. Figure 10 illustrates the scenarios for a few distinctive LPAs that are known to have tight or comparably relaxed planning constraints: Westminster and Newcastle upon Tyne were the most and least restrictive markets with respect to regulating office space (Cheshire and Hilber 2008). Reading and Darlington represent a relatively restrictive and a relatively relaxed LPA (Cheshire and Sheppard 1995). The predicted construction durations are in logarithms to improve comparability.

Our exercises suggest a substantial impact of supply constraints and developer market power on construction duration. Panel A of Table 6 suggests that, based on our baseline estimates, in 2015 the pure construction duration (i.e., the time from start to finish of a project *post initial planning approval*) in the 'average' LPA in England (with average housing demand shocks) would be 21 percent faster if the planning system were completely relaxed. Panel B then shows that reducing the restrictiveness by one standard deviation were to lead to a 9 percent reduction in construction duration. If we reduced both supply-side constraints and the developer market power by one standard deviation, the construction duration in the 'average' English LPA in 2015 would be 19 percent lower. Regarding the relative importance of these variables, both panels C and D suggest that reducing regulatory restrictiveness would have a larger quantitative impact compared with reducing other variables in the 'average' English LPA.

As Figure 10 illustrates, the impacts of regulatory constraints, physical (scarcity related) constraints, and developer market power vary significantly across locations. Physical constraints matter most in the densely developed borough of Westminster, while regulatory constraints are most important in the prosperous town of Reading. In Newcastle and Darlington, supply constraints have a relatively small positive impact on the construction duration. Local developer market power is more important in these locations.

4.5. Additional Results and Robustness Checks

In this section, we carry out several additional exercises and robustness checks.

Sites with different sizes: Small, medium and large

Construction sites with different sizes tend to have different building structures and are constructed by different types of developers. To study the heterogeneous effects of supply constraints and developer monopoly power on construction lags for sites with different sizes, we split our baseline estimation sample into three sub-samples (small sites, medium sites, and large sites) based on the number of dwellings within each site. We re-estimate equation (2) using these 3 sub-samples separately. Our OLS and IV estimates are reported in Appendix Table A3. The more credible IV estimates in columns (2), (4) and (6) suggest the following. First, while local demand shocks measured by ACLE speed up the build-out rates of small sites, they do not have a significant impact on the construction duration of either medium or large sites. The finding might be driven by the fact that large sites tend to have a pre-determined plan for build out and are thus not much influenced by short-term fluctuations in demand. Second, both regulatory constraint- and physical constraint-ACLE interactions have positive and statistically significant impacts on the construction duration across all three subsamples. This finding is in line with our baseline estimates and indicates that the impact of supply constraints on construction duration is consistent and substantial regardless of site size. Third, while we find a positive and significant impact of developer market power-ACLE interaction on construction duration for small and medium sites, this effect becomes insignificant for large sites. This is plausible as developers of large sites are less likely to be influenced by local competitiveness when making construction plans compared with the smaller ones.

Public vs. private sites

This section explores the extent to which the impacts of supply constraints and developer market power on construction lags varies between public and private sites. These two types of sites are developed by companies with different aims: private sites are constructed by profit-maximizing developers, while public dwellings are usually developed by housing associations and local authorities who aim to improve a local community's housing affordability. To explore this, we split our baseline estimation sites into two categories, public and private ones,[69] based on the share of public dwellings within each site. We then estimate these two subsamples separately and results are reported in Appendix Table A4. The findings of our preferred IV estimates, reported in columns (2) and (4), can be summarised as follows. First, while the ACLE has a negative and statistically significant impact on private units' construction duration, this impact is insignificant for public units. This is plausible as public units are provided by non-profit companies and the development process is not driven by local housing market conditions and firm profitability. Housing associations usually have 'pre-let' tenants to fill completed dwellings and thus are not likely to be influenced by the short-term fluctuation in housing demand. Second, in line with our baseline estimates, regulatory constraint-ACLE interaction has a positive and significant impact on construction duration, regardless of whether the site consists of predominately private or public units. Third, our estimate of the impact of local market power-ACLE interaction on construction duration is positive and statistically significant for private sites but it is insignificant for public sites. We interpret this finding as evidence showing that public dwellings are provided by developers with different incentives compared to private, profit maximising, developers. Local developer competitiveness is arguably not considered by housing associations when they provide public housing to the local community.

Sites with different housing mixes

To explore whether the mix of housing types on a site matters for the impact of our variables of interest, we split our baseline estimation sites into two categories: single-family house-type sites and flat-type sites,[70] based on the share of flats within each site. We then estimate these two subsamples separately and our estimates are reported in Appendix Table A5. The more rigorous columns (2) and (4) indicate that in line with our baseline results, the estimated coefficients for our variables of interest are positive and statistically significant, suggesting that supply constraints- and local market

[69] Private sites refer to sites with less than 50% public dwellings. Public sites refer to sites with more than or equal to 50% public dwellings.
[70] House-type sites refer to sites with less than 50% flats. Flat-type sites refer to sites with more than or equal to 50% flats.

power-ACLE interactions have positive impacts on construction duration regardless of the site's housing mix. However, there is a quantitative difference between the flat-type estimate and the single-family house-type estimate. We find that flat-type sites are more responsive to local demand shocks compared with single-family house-type sites, as the estimated coefficient for ACLE is more negative in column (4) compared with column (2). This finding is plausible as flat-type sites are more likely to be located in city centres and purchased by investors and young professionals, who are more sensitive to housing market conditions. Conversely, single-family house-type sites are more likely to be purchased by households with children, who tend to be more stable in residence. In addition, the impact of regulatory constraints is more pronounced for flat-type sites compared with single-family house-type sites, potentially because flat-type sites have more complex structures and are associated with more negative externalities. Therefore, existing homeowners have stronger incentives to oppose these projects and will put in more efforts to persuade local authorities to delay construction.

Plot-level results

In our baseline estimation sample, some construction sites may experience a change of developer during their build-out periods, and one might argue that this issue might not be fully controlled for since we could only include one developer fixed effect at the site level. To address this concern and to test for the robustness of our baseline results, we estimate a specification similar to equation (2) but at the dwelling level, as there is no change of developer occurring at the dwelling level. Our estimated results using over 2 million dwellings are reported in Appendix Table A6. The estimated coefficients for our variables of interest are all in line with the baseline findings and are robust across different specifications, suggesting that the main results are statistically consistent regardless of the level of observation used in our empirical analysis. In addition, Appendix Table A6 also presents the estimates for the dwelling-level control variables. Both columns (2) and (4) suggest that private dwellings and dwellings with more bedrooms have a longer construction duration. The estimated coefficient for the flat dummy is positive and statistically significant, indicating that it takes longer, all else equal, to build flats than single-family houses. This may be because flat-type sites usually require more complex structures.

Robustness check: Selection of instrumental variables

In our baseline specification, we jointly employ 3 separate instrumental variables to identify the refusal rate: the share of greenbelt land in 1973, the change in the delay rate, and the vote share of the Labour party in the 1983 General Election. One might be concerned that some of these instrumental variables may not be valid. In Appendix Table A7, we therefore report results for six alterations of our most rigorous baseline specification (column (6) of Table 2). The first three models drop one instrument at a time. Columns (4) to (6) then report estimates keeping only one of the three instruments at a time. The coefficients of interest remain stable across all six specifications, and the Kleibergen-Paap F-statistic suggests that weakness of identification is in general not a concern.

Robustness check: Alternative measures of developer competitiveness

We carry out two exercises to test for the robustness of our developer competitiveness estimates. At the site level, we draw a 10-km-radius buffer for each site and don't adjust the radius of the buffer based on LPA-level population density. We then compute each site's local market share using the constant 10-km-radius buffer and re-estimate equation (2) with this new measure of local market power. We report the findings in Appendix Table A8. The estimated coefficients of the interaction between local market share and ACLE are all positive and statistically significant. The estimates from columns (1) to (6) are also quantitatively similar to our baseline estimates in Table 2. In addition, at the LPA level, we use the top 5 largest developers' market share within each LPA instead of the Herfindahl-Hirschman index to measure developer competitiveness. We re-estimate equation (3) with this new measure and the results are reported in Appendix Table A9. In line with Table 4, the estimated coefficients for the interaction term between the top 5 largest developers' market share and LLDS are all positive and statistically significant across all four specifications, suggesting that developer market power has a positive impact on the construction duration-LLDS elasticity.

Robustness check: Alternative estimation period (between 1986 and 2018)

Next, we conduct an exercise to test for the robustness of our main findings with respect to the sample period. To do so, we re-estimate our baseline specification, however, we extend the sample period to the window '1986 to 2018', our baseline specification having focused on the shorter window between 1996 and 2015. This was because the NHBC records comparably fewer observations before 1996 and there might be a selection issue for sites recorded between 2016 and 2018. The results for the extended window are reported in Appendix Table A10. In line with Table 2, all the variables of interest have the expected signs and are statistically significant. These findings further support our baseline results and indicate that the impacts of supply constraints and developer market power are substantial in an even longer period of over 30 years.

Robustness check: Include sites with construction duration fewer than 70 days

We also conduct an exercise to test for the robustness of our main findings with respect to sample selection. We include sites with construction duration fewer than 70 days and re-estimate equation (2). The estimation results using this extended sample are reported in Appendix Table A11. The more credible IV estimates between columns (4) and (6) suggest that all the variables of interest have the expected signs and are statistically significant. These findings are in line with our baseline results in Table 2.

Robustness check: Controlling for the impact of Section 106 agreement-intensity

Finally, we explore the impact of the intensity of Section 106 agreements on site build-out rates. As noted in Section 2.1 Section 106 agreements are essentially a mechanism for extracting payment in kind whereby developers are granted the right to develop in return for providing a contribution for public consumption. This is usually in the form of a proportion of so-called ‘affordable housing’ (i.e., at a subsidized rate) to lower income households but historically has taken other forms such as land for parks, rights of way or contributions to local infrastructure. Section 106 agreements are negotiated between developers and LPAs. These negotiations are often complex and lengthy, and, moreover are often re-visited during the process of site construction as developers attempt to secure more favourable deals. S106-intensity can, therefore, be interpreted as another proxy measure for a different form of local ‘regulatory cost’.

We conduct this exercise by including an additional interaction term between ALDS and LPA-level S106 intensity into our baseline specification. We report results in Appendix Table A12. Columns (2) to (6) all suggest a positive and statistically significant impact of S106 agreement intensity-ACLE interaction on construction duration. This is in line with our expectation as Section 106 generates additional regulatory constraints, costs and possible delays for developers and thus further delays the site build-out. Appendix Table A12 also documents that our variables of interest – including the refusal rate interaction – remain stable and statistically significant, even after controlling for Section 106 intensity.

5. Conclusion

In this paper, we offer a rare insight into the determinants of 'construction lags' or the determinants of the rate at which construction sites are built out. We do this by analysing a unique and comprehensive dataset covering most residential developments in England between 1996 and 2015. We find that positive local demand shocks increase site build-out rates, but this impact is less for sites located in places with severe housing supply constraints and for sites constructed by developers with more market power. Our main results are consistent across different specifications and robustness checks.

Our paper raises several interesting questions for future research. We have discussed the impact of the S106 agreements on site build-out rate in section 4.5, yet further investigation could be undertaken to explore the underlying mechanisms. Besides, Table A1 suggests that top companies have high market shares in England. It is also an intriguing and important question to explore the potential causes of monopolistic power in the British housebuilding industry and discuss whether a more restrictive planning system leads to this market concentration.

Tables

Table 1: Descriptive Statistics (Baseline Sample)

	Observations	Mean	SD	Max	Min
Panel A: Construction site level					
Construction duration (days)	119003	594.04	493.29	7720	71
Number of dwellings per site	119003	18.38	42.18	1075	1
Average number of bedrooms	115882	3.25	1.29	100	0
Share of flats	119003	0.16	0.35	1	0
Change of developer dummy	119003	0.01	0.11	1	0
Public project ratio	119003	0.06	0.22	1	0
Local market share	119003	0.07	0.14	1	0.0002
Distance to CBD (km)	119003	11.94	6.64	45.75	0.03
Average temperature (Celsius)	118269	10.23	0.62	11.83	7.22
Average wind speed	118269	3.98	0.7	7.93	2.08
Panel B: Local planning authority level, panel data (N=353, T=20)					
Log (predicted employment)	6978	10.79	0.64	13.12	8.33
Construction duration (days)	6978	652.6	272.29	4326	141
Avg. number of dwellings	6978	24.48	26.98	370	1
Avg. number of bedrooms	6978	3.18	0.62	21.13	1
Share of flats	6978	0.17	0.20	1	0
Share of public units	6978	0.08	0.11	1	0
Panel C: Local planning authority level, cross-section (N=353)					
Avg. refusal rate	353	0.19	0.08	0.43	0.003
% Developable land developed	353	0.26	0.23	0.98	0.01
Herfindahl-Hirschman index	353	0.09	0.04	0.31	0.02
Top 5 developers' market share	353	0.55	0.11	0.84	0.26
Panel D: Local planning authority level, instrumental variables, cross-section (N=353)					
Change in delay rate	353	-0.03	0.22	0.53	-0.63
Share of votes for Labour	353	0.47	0.31	1	0
Share of greenbelt land in 1973	353	0.09	0.22	1	0
Population density in 1911	353	733.27	2561.63	22028.8	3.25

Table 2: Baseline Estimation Results

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	OLS	IV	IV	IV
ACLE [1)]	-7.3747***	-3.7169***	-3.6106***	-7.4011***	-3.7428***	-3.6344***
	(0.6073)	(0.5935)	(0.5962)	(0.6123)	(0.6053)	(0.6073)
Refusal rate × ACLE	1.0657***	1.4625***	1.4588***	1.5798***	2.3682***	2.3207***
	(0.2617)	(0.2394)	(0.2416)	(0.4206)	(0.3844)	(0.3812)
% Developed × ACLE	0.6910**	0.4257	0.4444	1.0989***	1.1860***	1.1713***
	(0.2979)	(0.2802)	(0.2782)	(0.4168)	(0.4404)	(0.4351)
Local market share × ACLE	2.2937***	1.2535***	1.1572***	2.3799***	1.4011***	1.2987***
	(0.1881)	(0.1528)	(0.1532)	(0.1964)	(0.1655)	(0.1657)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Weather controls	No	No	Yes	No	No	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period	1996-2015					
N	115882	102163	101452	115882	102163	101452
R^2	0.3433	0.5996	0.6015			
Kleibergen-Paap F				24.26	24.91	25.08

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 3: First Stage Results

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Refusal rate ×ACLE	%Developed ×ACLE	Refusal rate ×ACLE	%Developed ×ACLE	Refusal rate ×ACLE	%Developed ×ACLE
ACLE [1)]	0.0242	0.0309	0.0097	0.0308	0.0087	0.0298
	(0.0530)	(0.0425)	(0.0550)	(0.0437)	(0.0551)	(0.0435)
Change in delay rate × ACLE	-0.1920***	-0.0554	-0.1866***	-0.0519	-0.1865***	-0.0510
	(0.0660)	(0.0381)	(0.0667)	(0.0363)	(0.0668)	(0.0357)
% Labour vote in 1983 × ACLE	-0.5169***	0.5328***	-0.5235***	0.5360***	-0.5232***	0.5360***
	(0.0576)	(0.0463)	(0.0563)	(0.0468)	(0.0564)	(0.0466)
% Greenbelt in 1973 × ACLE	0.3102***	0.0841***	0.3104***	0.0790***	0.3115***	0.0801***
	(0.0525)	(0.0269)	(0.0531)	(0.0265)	(0.0530)	(0.0263)
Pop. density in 1911 × ACLE	0.1564**	0.4331***	0.1574**	0.4271***	0.1577**	0.4272***
	(0.0705)	(0.0594)	(0.0705)	(0.0611)	(0.0704)	(0.0610)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	No	Yes	Yes	Yes	Yes
Weather controls	No	No	No	No	Yes	Yes
Soil conditions	No	No	No	No	Yes	Yes
Time Period			1996-2015			
N	115882	115882	102163	102163	101452	101452
R^2	0.7182	0.8095	0.7684	0.8431	0.7688	0.8437

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 4: Estimation Results at the LPA Level

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	IV	IV
LLDS [1)]	-0.3598	-0.4896**	-0.5732**	-0.7081***
	(0.2513)	(0.2138)	(0.2793)	(0.2414)
Refusal rate × LLDS	0.1179	0.2140**	0.5009**	0.6662***
	(0.1015)	(0.0843)	(0.2402)	(0.1969)
% Developed × LLDS	0.1853**	0.2183***	0.5186***	0.4471***
	(0.0881)	(0.0750)	(0.1232)	(0.1255)
Herfindahl-Hirschman Index × LLDS	0.2630*	0.2107*	0.4425**	0.4193***
	(0.1412)	(0.1094)	(0.1735)	(0.1391)
Log (number of dwellings)		0.2227***		0.2235***
		(0.0064)		(0.0065)
Avg. number of bedrooms per dwelling		0.0446**		0.0437**
		(0.0179)		(0.0179)
Share of flats		-0.1302***		-0.1418***
		(0.0433)		(0.0433)
Share of public units		-0.4560***		-0.4500***
		(0.0464)		(0.0461)
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Time Period		1996-2015		
Number of LPAs		353		
N	6978	6978	6978	6978
R^2	0.4052	0.6021		
Kleibergen-Paap F			15.28	15.33

Notes: [1)] LLDS refers to the natural logarithm of predicted local employment. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 5: First Stage Results at the LPA Level

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Refusal rate ×LLDS	%Developed ×LLDS	Refusal rate ×ACLE	%Developed ×LLDS
LLDS [1)]	0.1926**	0.2290***	0.1914**	0.2298***
	(0.0852)	(0.0836)	(0.0851)	(0.0837)
Change in delay rate × LLDS	-0.0970*	-0.0498	-0.0961*	-0.0491
	(0.0555)	(0.0463)	(0.0555)	(0.0461)
Share Labour vote in 1983 × LLDS	-0.3306***	0.5071***	-0.3333***	0.5051***
	(0.0561)	(0.0495)	(0.0563)	(0.0497)
Share greenbelt in 1973 × LLDS	0.3221***	0.0921***	0.3205***	0.0903***
	(0.0514)	(0.0295)	(0.0513)	(0.0295)
Population density in 1911 × LLDS	0.0066	0.4600***	0.0064	0.4601***
	(0.0563)	(0.0503)	(0.0564)	(0.0501)
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Other controls	No	No	Yes	Yes
Time Period	1996-2015			
N	6978	6978	6978	6978
R^2	1	1	1	1

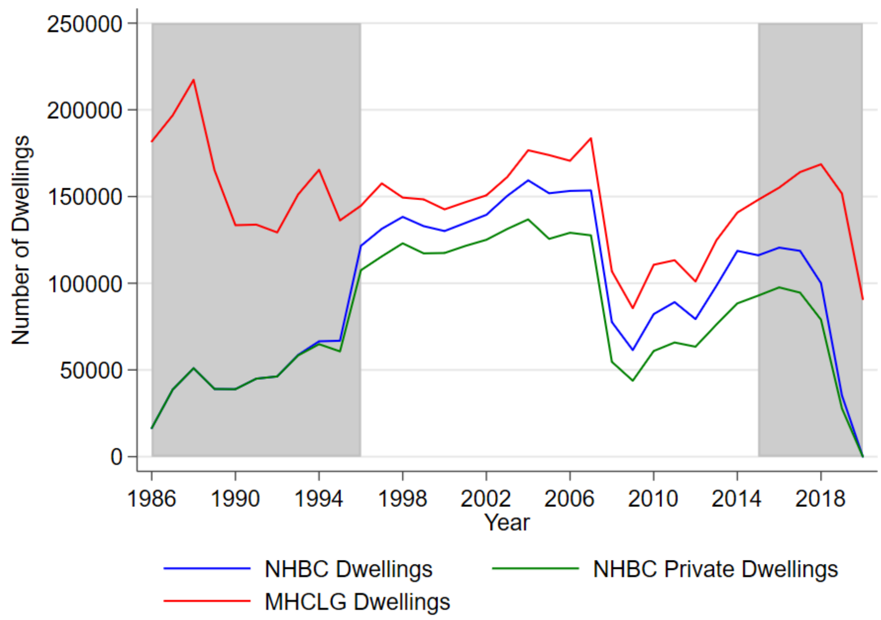
Notes: [1)] LLDS refers to the natural logarithm of predicted local employment. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 6: Effect of Shifts in LLDS on Construction Duration in Average English LPA (Counterfactual Outcomes)

Variable	Value in 1996	Value in 2015	SD	Max	Min
Predicted construction duration (in days)	559	744	222	1587	268
Panel A: supply constraints, market power, and demand shock set to zero sequentially					
Predicted without any planning refusals	559	588	175	1253	211
- and share developed set to zero	559	547	163	1166	197
- and HHI set to zero	559	471	141	1005	170
- and independent effect of LLDS removed	559	521	156	1111	188
Panel B: supply constraints and market power lowered by 1 SD sequentially					
Predicted with refusal rate lowered by 1 SD	559	677	202	1444	244
- and share developed lowered by 1 SD	559	635	190	1355	229
- and HHI lowered by 1 SD	559	599	179	1276	215
Panel C: supply constraints and market power set to zero separately					
Predicted without any planning refusals	559	588	175	1253	211
Predicted with share developed set to zero	559	693	207	1477	249
Predicted with HHI set to zero	559	642	192	1368	231
Panel D: supply constraints and market power lowered by 1 SD separately					
Predicted with refusal rate lowered by 1 SD	559	677	202	1444	244
Predicted with share developed lowered by 1 SD	559	699	209	1489	251
Predicted with HHI lowered by 1 SD	559	701	209	1495	252

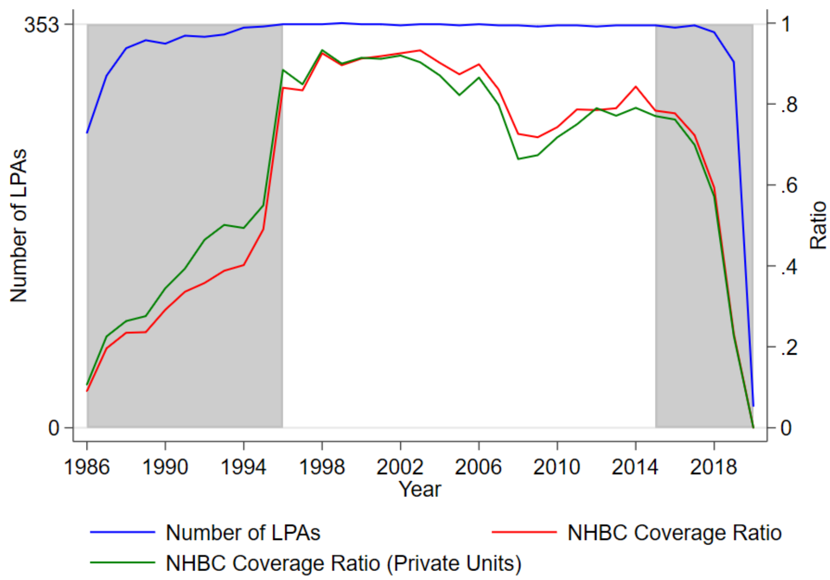
Figures

Fig. 1: NHBC and MHCLG Number of Dwellings



Note: The shaded years (1986 to 1995, 2016 to 2020) are excluded from the baseline estimation sample. The major estimation period is between 1996 and 2015 (20 years).

Fig. 2: NHBC Coverage Ratio



Note: NHBC coverage ratio represents the ratio of NHBC units relative to the units recorded by MHCLG. Number of LPAs denotes the number of local authorities in England that are covered in the NHBC dataset.

Fig. 3: Histograms

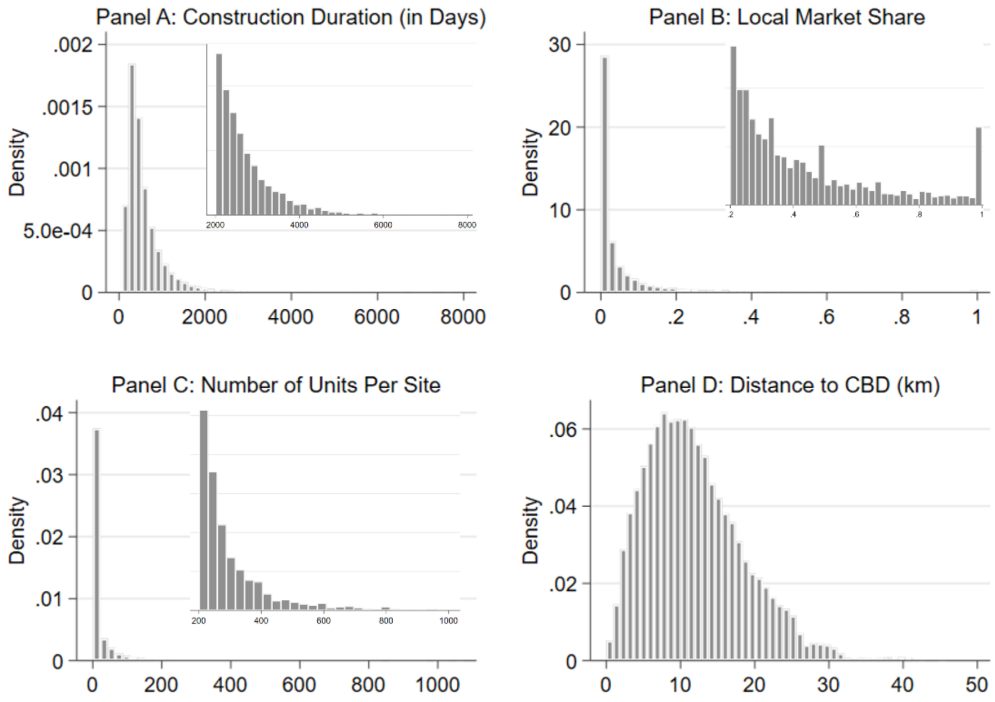
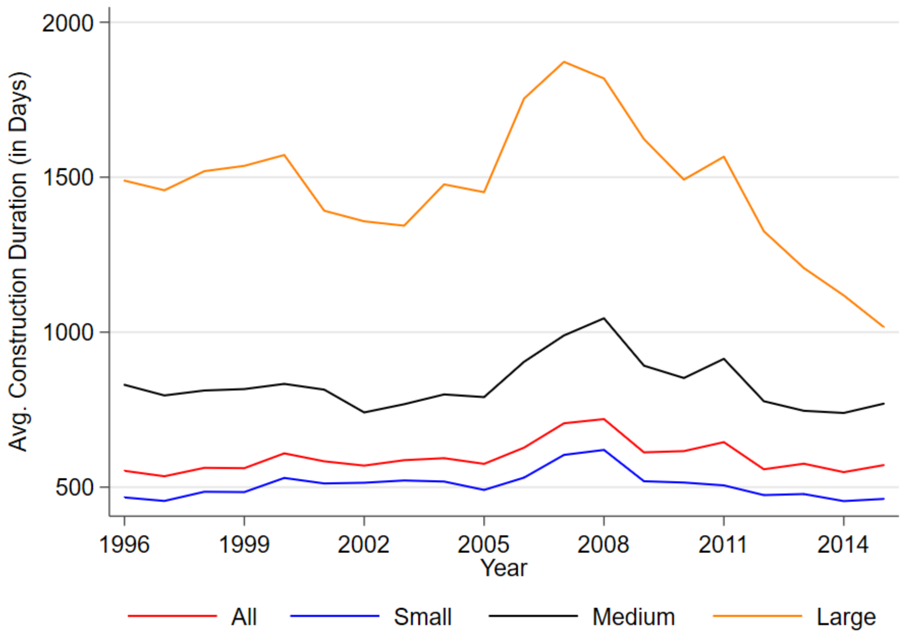
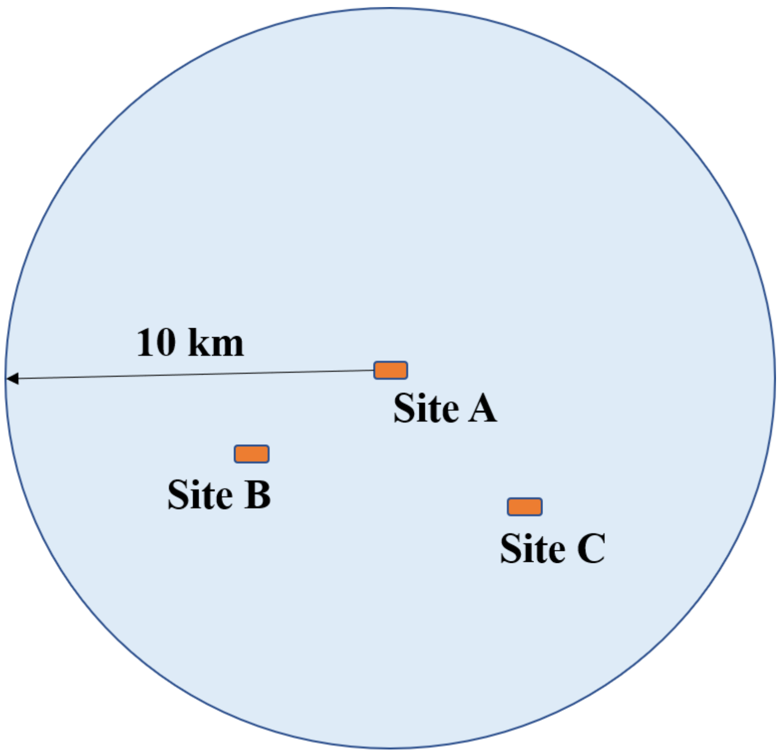


Fig. 4: Average Construction Duration in England



Note: Small refers to projects with less than 25 units. Medium refers to projects with 25 to 100 unit. Large refers to projects with more than 100 units (Ball, 2011).

Fig. 5: Measure Local Market Share



Note: The radius of the circle is normalized based on TTWA population density in 2001.

Fig. 6: Market Share and Herfindahl-Hirschman Index in England

Fig. 7: Average Size of Construction Site in England

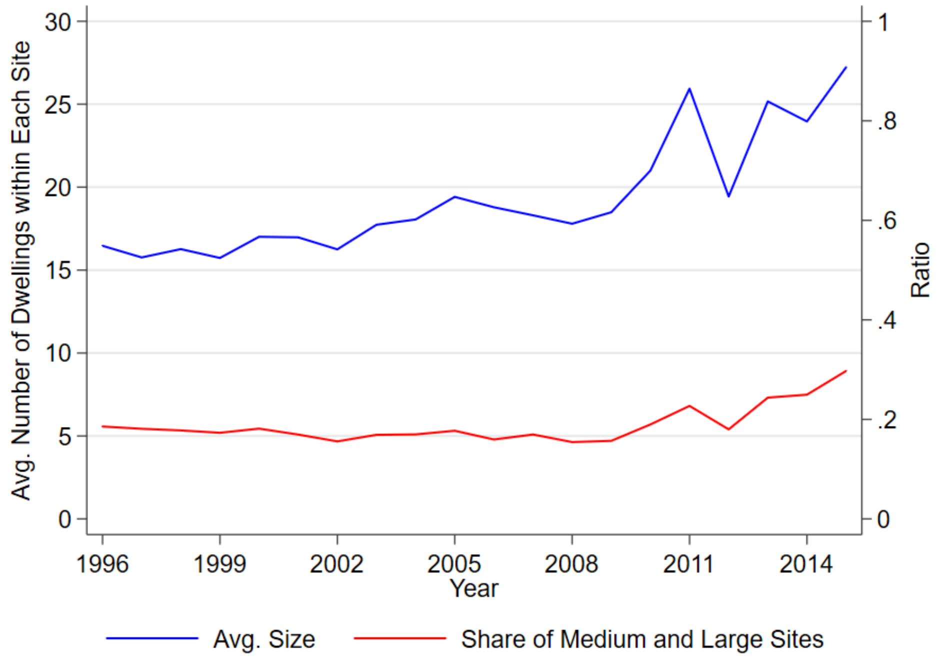


Fig. 8: Impact of Removing Supply Constraints and Developer Market Power on Construction Duration in Average English LPA

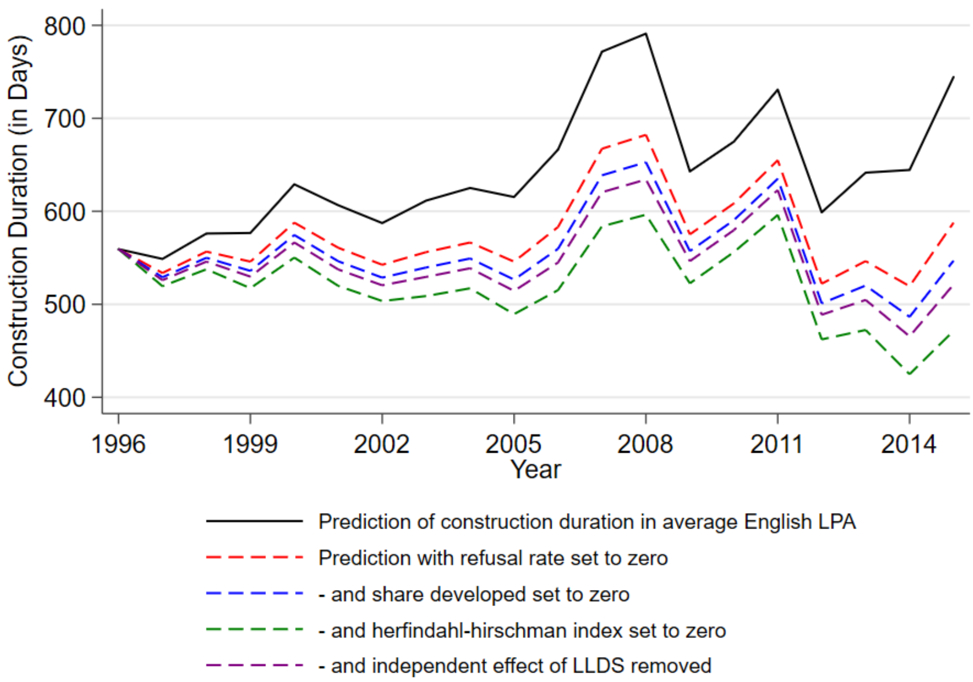


Fig. 9: Impact of Reducing Supply Constraints and Developer Market Power on Construction Duration in Average English LPA

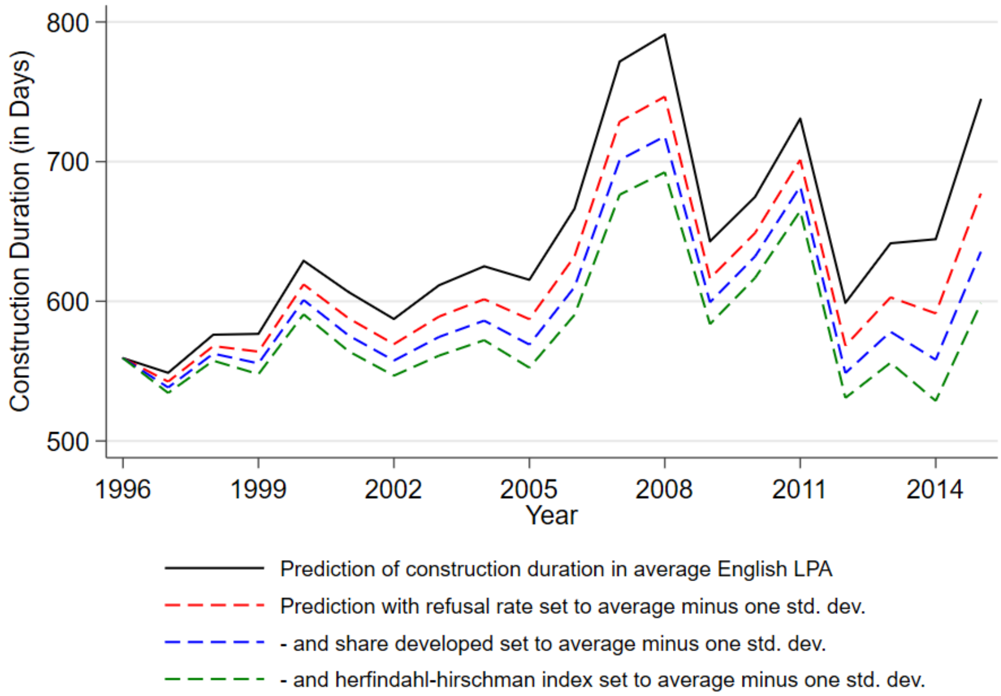
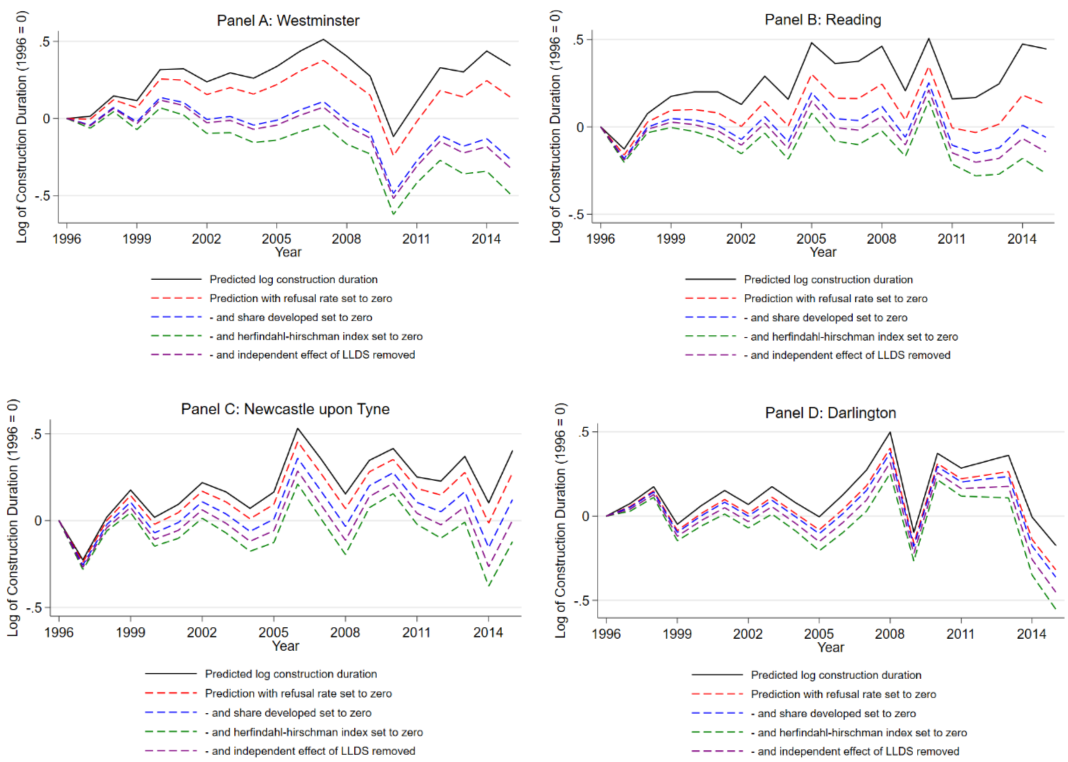


Fig. 10: Predicted Log of Construction Duration in Selected LPAs



Appendices

Appendix A: Appendix Tables

Table A1: Top 10 Developers' Market Share in England Between 1996 and 2015

Rank	Developer name	Market share
1	Taylor Wimpey UK Limited	12.5%
2	BDW Trading Limited (Barratt)	11.8%
3	Persimmon Homes Ltd	8.3%
4	Bellway Homes Ltd	4.7%
5	Redrow Homes Ltd	2.6%
6	Galliford Try Plc	2.2%
7	Berkeley Group Plc	2%
8	Crest Nicholson Residential Limited	1.9%
9	Bovis Homes Ltd	1.8%
10	Westbury Homes (Holdings) Ltd	1.6%

Note: Top 10 developers account for 49.3% of all new dwellings in our estimation sample.

Table A2: Baseline Results with Controls

Specifications Dependent variable:	(1) Log(lag) OLS	(2) Log(lag) OLS	(3) Log(lag) OLS	(4) Log(lag) IV	(5) Log(lag) IV	(6) Log(lag) IV
ACLE [1)]	-7.3747***	-3.7169***	-3.6106***	-7.4011***	-3.7428***	-3.6344***
	(0.6073)	(0.5935)	(0.5962)	(0.6123)	(0.6053)	(0.6073)
Refusal rate × ACLE	1.0657***	1.4625***	1.4588***	1.5798***	2.3682***	2.3207***
	(0.2617)	(0.2394)	(0.2416)	(0.4206)	(0.3844)	(0.3812)
% Developed × ACLE	0.6910**	0.4257	0.4444	1.0989***	1.1860***	1.1713***
	(0.2979)	(0.2802)	(0.2782)	(0.4168)	(0.4404)	(0.4351)
Local market share × ACLE	2.2937***	1.2535***	1.1572***	2.3799***	1.4011***	1.2987***
	(0.1881)	(0.1528)	(0.1532)	(0.1964)	(0.1655)	(0.1657)
Log (number of dwellings)	0.2398***	0.3657***	0.3674***	0.2396***	0.3653***	0.3670***
	(0.0025)	(0.0046)	(0.0045)	(0.0025)	(0.0047)	(0.0045)
Avg. # of bedrooms	0.0717***	0.0811***	0.0797***	0.0717***	0.0810***	0.0797***
	(0.0088)	(0.0113)	(0.0112)	(0.0087)	(0.0113)	(0.0112)
Share of 'flats'	-0.1559***	-0.1569***	-0.1533***	-0.1559***	-0.1568***	-0.1532***
	(0.0158)	(0.0186)	(0.0183)	(0.0158)	(0.0186)	(0.0183)
Change of developer dummy	0.6131***	0.5513***	0.5496***	0.6129***	0.5509***	0.5492***
	(0.0198)	(0.0220)	(0.0219)	(0.0198)	(0.0220)	(0.0219)
Public project ratio	-0.4544***	-0.3737***	-0.3725***	-0.4545***	-0.3733***	-0.3721***
	(0.0154)	(0.0169)	(0.0169)	(0.0154)	(0.0170)	(0.0170)
Log (distance to CBD)	0.0222***	0.0212***	0.0148***	0.0221***	0.0210***	0.0146***
	(0.0045)	(0.0045)	(0.0043)	(0.0045)	(0.0045)	(0.0043)
Weather – temperature			-0.7863***			-0.7834***
			(0.1025)			(0.1025)
Weather – wind speed			0.0539**			0.0539**
			(0.0266)			(0.0266)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period			1996-2015			
N	115882	102163	101452	115882	102163	101452
R^2	0.3433	0.5996	0.6015			
Kleibergen-Paap F				24.26	24.91	25.08

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A3: Different-size Sites

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	IV	OLS	IV	OLS	IV
Project:	Small	Small	Medium	Medium	Large	Large
ACLE [1)]	-4.9539***	-4.9211***	-1.4448	-1.2145	3.2196	3.2936
	(0.7422)	(0.7522)	(1.2784)	(1.3220)	(2.0714)	(2.1232)
Refusal rate × ACLE	1.1235***	1.6009***	2.0912***	3.5126***	0.6672	3.3720**
	(0.2671)	(0.3938)	(0.5150)	(0.8873)	(0.8325)	(1.5777)
% Developed × ACLE	0.3714	0.9424*	1.0164**	1.6898***	0.1798	1.3521*
	(0.3345)	(0.5251)	(0.4059)	(0.5075)	(0.6417)	(0.7530)
Local market share × ACLE	0.7230***	0.8387***	0.4964*	0.5902**	0.0064	0.1576
	(0.2064)	(0.2224)	(0.2658)	(0.2736)	(0.4034)	(0.4110)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	Yes	Yes	Yes	Yes	Yes	Yes
Weather controls	Yes	Yes	Yes	Yes	Yes	Yes
Soil conditions	Yes	Yes	Yes	Yes	Yes	Yes
Time Period			1996-2015			
N	80271	80271	15656	15656	4620	4620
R^2	0.5075		0.5064		0.5949	
Kleibergen-Paap F		22.33		18.31		13.58

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A4: Public/Private Sites

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	IV	OLS	IV
Project:	Private	Private	Public	Public
ACLE [1)]	-3.5012***	-3.5071***	-0.2275	-0.0095
	(0.6124)	(0.6218)	(2.1382)	(2.1439)
Refusal rate × ACLE	1.5728***	2.3274***	1.8123***	4.0183***
	(0.2406)	(0.3869)	(0.6633)	(1.4223)
% Developed × ACLE	0.3327	1.0287**	0.3196	0.8108
	(0.2928)	(0.4665)	(0.5423)	(0.6931)
Local market share × ACLE	1.2036***	1.3343***	0.2317	0.3387
	(0.1614)	(0.1750)	(0.3625)	(0.3691)
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes
Developer FEs	Yes	Yes	Yes	Yes
Weather controls	Yes	Yes	Yes	Yes
Soil conditions	Yes	Yes	Yes	Yes
Time Period	1996-2015			
N	95535	95535	5461	5461
R^2	0.6086		0.6752	
Kleibergen-Paap F		24.13		11.29

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A5: Sites with Different Structures

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	IV	OLS	IV
Project:	Houses	Houses	Flats	Flats
ACLE [1]	-2.6104***	-2.5319***	-8.0853***	-9.4829***
	(0.6063)	(0.6134)	(1.6562)	(1.7683)
Refusal rate × ACLE	1.5575***	1.9803***	1.0138*	3.6109**
	(0.2468)	(0.3562)	(0.5178)	(1.4283)
% Developed × ACLE	0.4227	0.9728*	0.8551	1.9947**
	(0.2985)	(0.5133)	(0.5374)	(0.9238)
Local market share × ACLE	1.0643***	1.1454***	2.0057***	2.4192***
	(0.1598)	(0.1721)	(0.4547)	(0.4980)
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes
Seasonality [2]	Yes	Yes	Yes	Yes
Developer FEs	Yes	Yes	Yes	Yes
Weather controls	Yes	Yes	Yes	Yes
Soil conditions	Yes	Yes	Yes	Yes
Time Period	1996-2015			
N	83701	83701	15012	15012
R^2	0.6274		0.5733	
Kleibergen-Paap F		20.07		8.51

Notes: [1] ACLE refers to the annualised change in local employment. [2] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A6: Estimation at the Dwelling Level

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	IV	IV
ACLE [1)]	-6.4516***	-6.4439***	-5.5913***	-5.6598***
	(0.8877)	(0.8689)	(0.9030)	(0.8819)
Refusal rate × ACLE	0.7978*	0.7225	2.9350***	2.6813***
	(0.4426)	(0.4408)	(0.9140)	(0.8828)
% Developed × ACLE	0.8604*	0.7965*	2.0224***	1.8201***
	(0.4969)	(0.4713)	(0.6659)	(0.6374)
Local market share × ACLE	1.8897***	1.8566***	1.9075***	1.8674***
	(0.2782)	(0.2722)	(0.2837)	(0.2766)
Number of bedrooms		0.0059***		0.0059***
		(0.0015)		(0.0016)
Dummy – flat		0.1388***		0.1389***
		(0.0052)		(0.0052)
Dummy – public dwelling		-0.0667***		-0.0659***
		(0.0065)		(0.0065)
Year FEs	Yes	Yes	Yes	Yes
Construction site FEs	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes
Time Period	1996-2015			
N	2251007	2234289	2251007	2234289
R^2	0.6347	0.6393		
Kleibergen-Paap F			18.73	18.84

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A7: Selection of Instrumental Variables

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	IV exclude change in delay rate	IV exclude share of labour party	IV exclude greenbelt	greenbelt only	change in delay rate only	share of labour party only
ACLE [1]	-3.6358***	-3.5945***	-3.6463***	-3.5881***	-3.6054***	-3.6576***
	(0.6079)	(0.6081)	(0.6109)	(0.6087)	(0.6109)	(0.6160)
Refusal rate × ACLE	2.3499***	2.0719***	2.5753***	2.0074***	2.2094**	2.8200***
	(0.4100)	(0.4564)	(0.6384)	(0.4940)	(0.8787)	(0.8546)
% Developed × ACLE	1.1892***	1.3791**	1.3252**	1.3824***	1.4051**	1.4740**
	(0.4493)	(0.5329)	(0.5390)	(0.5295)	(0.5549)	(0.6584)
Local market share × ACLE	1.3029***	1.2919***	1.3355***	1.2860***	1.3071***	1.3710***
	(0.1674)	(0.1644)	(0.1822)	(0.1651)	(0.1842)	(0.2013)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	Yes	Yes	Yes	Yes	Yes	Yes
Weather controls	Yes	Yes	Yes	Yes	Yes	Yes
Soil conditions	Yes	Yes	Yes	Yes	Yes	Yes
Time Period	1996-2015					
N	101452	101452	101452	101452	101452	101452
Kleibergen-Paap F	31.85	13.51	16.63	15.25	7.74	17.01

Notes: [1] ACLE refers to the annualised change in local employment. [2] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A8: Different Measure of Local Competitiveness (10km buffer with no adjustment)

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	OLS	IV	IV	IV
ACLE [1)]	-7.5518***	-3.8078***	-3.6953***	-7.5829***	-3.8457***	-3.7307***
	(0.6081)	(0.5944)	(0.5972)	(0.6132)	(0.6064)	(0.6085)
Refusal rate × ACLE	1.0404***	1.4536***	1.4520***	1.5224***	2.3461***	2.3034***
	(0.2612)	(0.2385)	(0.2408)	(0.4180)	(0.3829)	(0.3800)
% Developed × ACLE	0.8377***	0.5299*	0.5435*	1.2322***	1.2959***	1.2769***
	(0.2974)	(0.2808)	(0.2789)	(0.4167)	(0.4445)	(0.4394)
Local market share × ACLE	1.8849***	1.0686***	0.9959***	1.9657***	1.2153***	1.1366***
	(0.1604)	(0.1314)	(0.1321)	(0.1697)	(0.1456)	(0.1461)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Weather controls	No	No	Yes	No	No	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period			1996-2015			
N	115882	102163	101452	115882	102163	101452
R^2	0.3432	0.5996	0.6015			
Kleibergen-Paap F				24.13	24.79	24.96

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A9: Estimation Results at the LPA Level (Top 5 Developer Market Share)

Specifications	(1)	(2)	(3)	(4)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	IV	IV
LLDS [1)]	-0.3559	-0.4855**	-0.5722**	-0.7113***
	(0.2513)	(0.2137)	(0.2796)	(0.2420)
Refusal rate × LLDS	0.1675	0.2384***	0.5724**	0.7329***
	(0.1027)	(0.0858)	(0.2577)	(0.2132)
% Developed × LLDS	0.1826**	0.2155***	0.5416***	0.4692***
	(0.0886)	(0.0752)	(0.1248)	(0.1297)
Top 5 developers’ market share × LLDS	0.3379***	0.2362**	0.5394***	0.4835***
	(0.1276)	(0.1051)	(0.1677)	(0.1399)
Log (number of dwellings)		0.2226***		0.2233***
		(0.0064)		(0.0065)
Average number of bedrooms per dwelling		0.0446**		0.0436**
		(0.0180)		(0.0180)
Share of flats		-0.1311***		-0.1444***
		(0.0433)		(0.0433)
Share of public units		-0.4565***		-0.4504***
		(0.0465)		(0.0462)
LPA FEs	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Time Period	1996-2015			
Number of LPAs	353			
N	6978	6978	6978	6978
R^2	0.4056	0.6022		
Kleibergen-Paap F			13.31	13.34

Notes: [1)] LLDS refers to the natural logarithm of predicted local employment. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A10: Baseline Specification with Extended Period

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	OLS	IV	IV	IV
ACLE [1)]	-1.7191***	-0.8622*	-0.7911	-1.7867***	-0.9440*	-0.8712
	(0.5639)	(0.5166)	(0.5199)	(0.5761)	(0.5400)	(0.5421)
Refusal rate × ACLE	0.9803***	1.1657***	1.1633***	1.6773***	1.9832***	1.9641***
	(0.2106)	(0.1822)	(0.1824)	(0.3727)	(0.3491)	(0.3469)
% Developed × ACLE	1.7420***	1.4596***	1.4592***	2.5921***	2.5289***	2.5080***
	(0.2638)	(0.2269)	(0.2255)	(0.3650)	(0.3305)	(0.3262)
Local market share × ACLE	1.6204***	0.7907***	0.7406***	1.7375***	0.9295***	0.8776***
	(0.1362)	(0.1146)	(0.1157)	(0.1448)	(0.1259)	(0.1267)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Weather controls	No	No	Yes	No	No	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period	1986-2018					
N	139432	124313	123507	139432	124313	123507
R^2	0.3405	0.5921	0.5938			
Kleibergen-Paap F				21.26	21	21.1

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A11: Baseline Specification Including Sites with Fewer Than 70 days Duration

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	OLS	IV	IV	IV
ACLE [1)]	-10.3163***	-4.5292***	-4.4202***	-10.3300***	-4.5409***	-4.4300***
	(0.6986)	(0.6526)	(0.6579)	(0.7052)	(0.6649)	(0.6696)
Refusal rate × ACLE	1.1463***	1.4910***	1.4882***	1.6528***	2.3642***	2.3229***
	(0.2797)	(0.2499)	(0.2520)	(0.4651)	(0.4153)	(0.4119)
% Developed × ACLE	0.6903**	0.3859	0.3940	1.2758***	1.3077***	1.2890***
	(0.3126)	(0.3070)	(0.3042)	(0.4136)	(0.4831)	(0.4765)
Local market share × ACLE	2.2389***	1.2404***	1.1388***	2.3398***	1.3974***	1.2907***
	(0.1820)	(0.1531)	(0.1549)	(0.1902)	(0.1670)	(0.1685)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Weather controls	No	No	Yes	No	No	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period	1996-2015					
N	116843	102805	102088	116843	102805	102088
R^2	0.3108	0.5724	0.5741			
Kleibergen-Paap F				24.28	24.95	25.12

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table A12: Estimation Results with S106 Intensity Variable

Specifications	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)	Log(lag)
	OLS	OLS	OLS	IV	IV	IV
ACLE [1)]	-7.4189***	-3.7972***	-3.6933***	-7.4478***	-3.8364***	-3.7295***
	(0.6040)	(0.5898)	(0.5925)	(0.6077)	(0.5994)	(0.6016)
S106 intensity × ACLE	0.3920	0.5892**	0.6002**	0.4950**	0.7668***	0.7679***
	(0.2476)	(0.2434)	(0.2406)	(0.2411)	(0.2289)	(0.2266)
Refusal rate × ACLE	1.1222***	1.5479***	1.5452***	1.5368***	2.3173***	2.2741***
	(0.2605)	(0.2386)	(0.2406)	(0.4039)	(0.3691)	(0.3662)
% Developed × ACLE	0.7566**	0.5192*	0.5394**	1.1369***	1.2500***	1.2398***
	(0.2952)	(0.2647)	(0.2619)	(0.3963)	(0.4142)	(0.4083)
Local market share × ACLE	2.3015***	1.2666***	1.1704***	2.3747***	1.3981***	1.2962***
	(0.1881)	(0.1529)	(0.1532)	(0.1957)	(0.1637)	(0.1638)
LPA FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
Site characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Seasonality [2)]	Yes	Yes	Yes	Yes	Yes	Yes
Developer FEs	No	Yes	Yes	No	Yes	Yes
Weather controls	No	No	Yes	No	No	Yes
Soil conditions	No	No	Yes	No	No	Yes
Time Period	1996-2015					
N	115882	102163	101452	115882	102163	101452
R^2	0.3434	0.5997	0.6016			
Kleibergen-Paap F				22.05	22.76	22.94

Notes: [1)] ACLE refers to the annualised change in local employment. [2)] We control for the seasonality of real estate development by including both the start month and the completion month FEs in the specification. Standard errors are clustered at the LPA level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Appendix B: Appendix Figures

Fig. B1: Average Refusal Rate

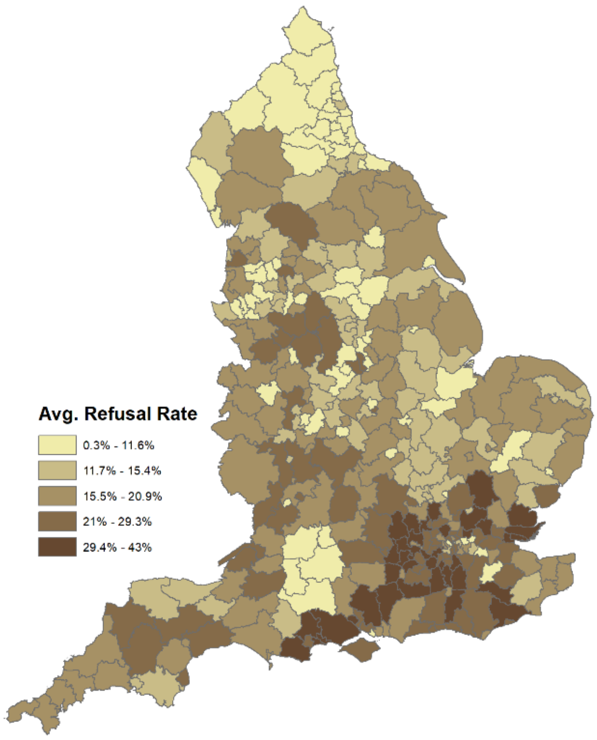


Fig. B2: S106 Intensity

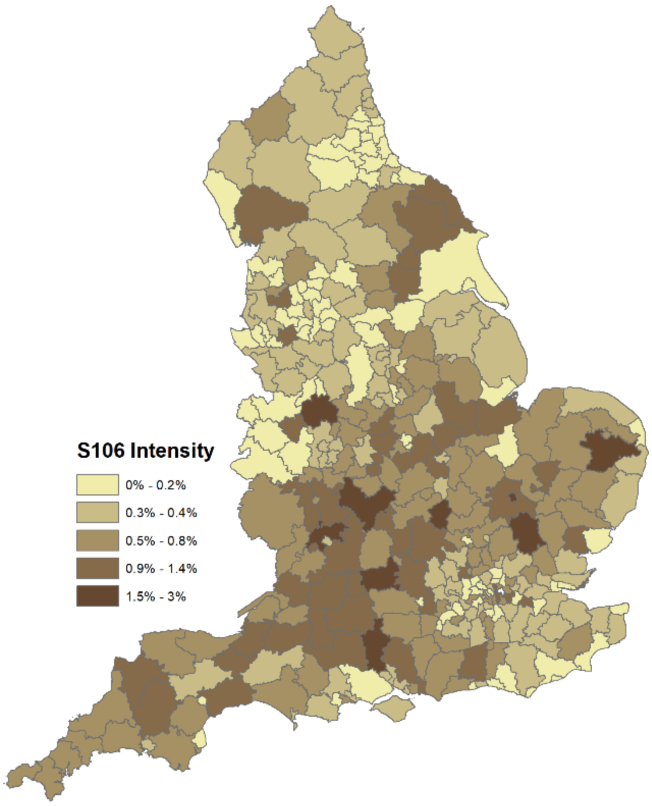


Fig. B3: Percentage Change in Predicted Employment

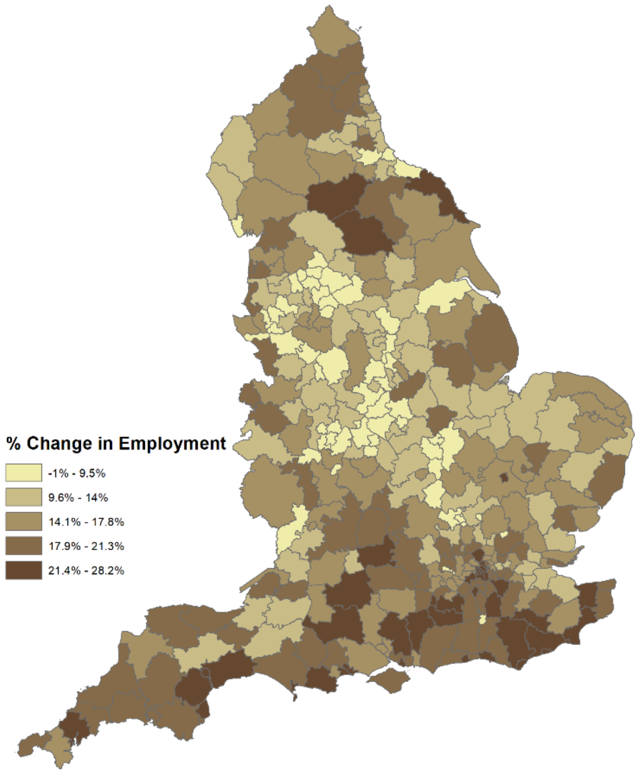


Fig. B4: Average Construction Duration (in Days) at the Site Level

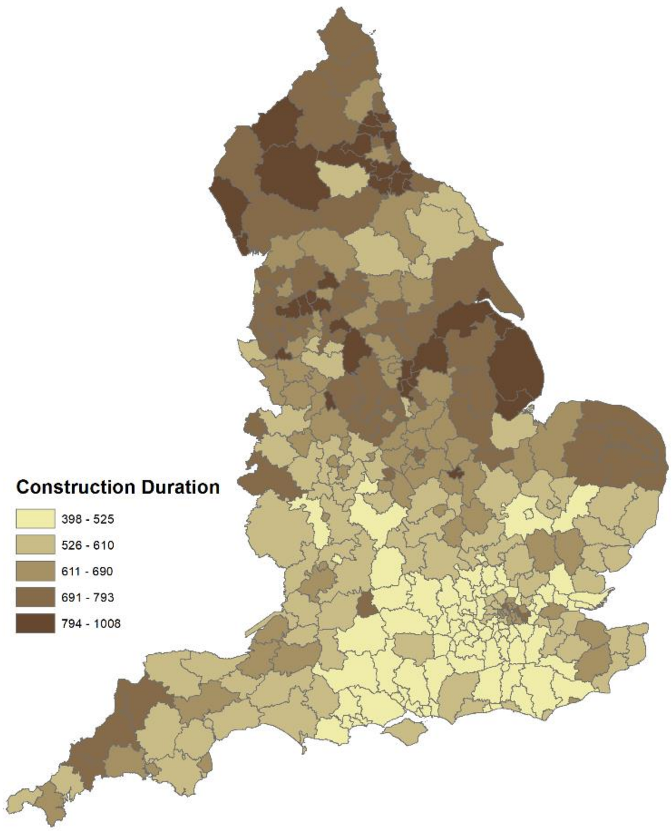
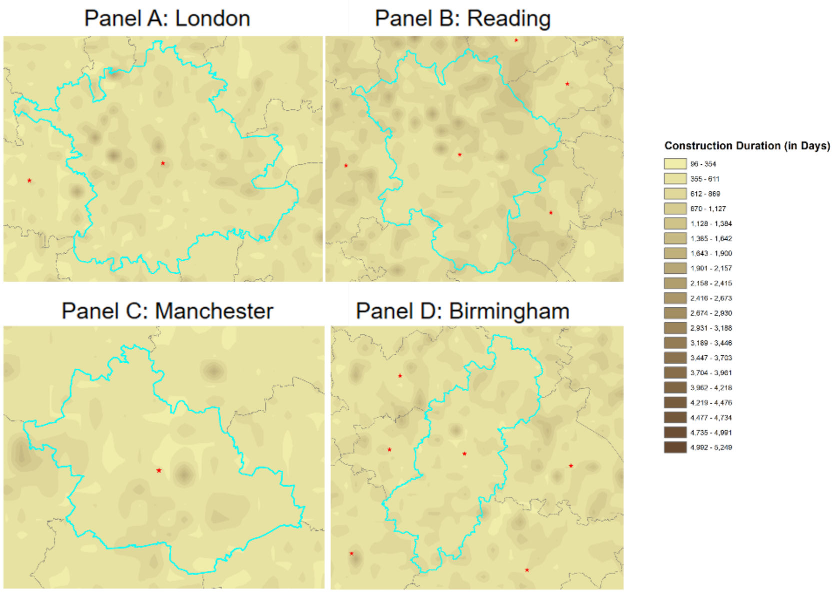


Fig. B5: Construction Duration Within Cities



Note: Panels A to D show the smoothed construction duration using geocoded sites and the inverse distance weighting technique. The red stars represent the centroids of each TTWA (as a proxy for city centre). In general, sites away from the city centre have a longer construction duration.

Fig. B6: Herfindahl-Hirschman Index

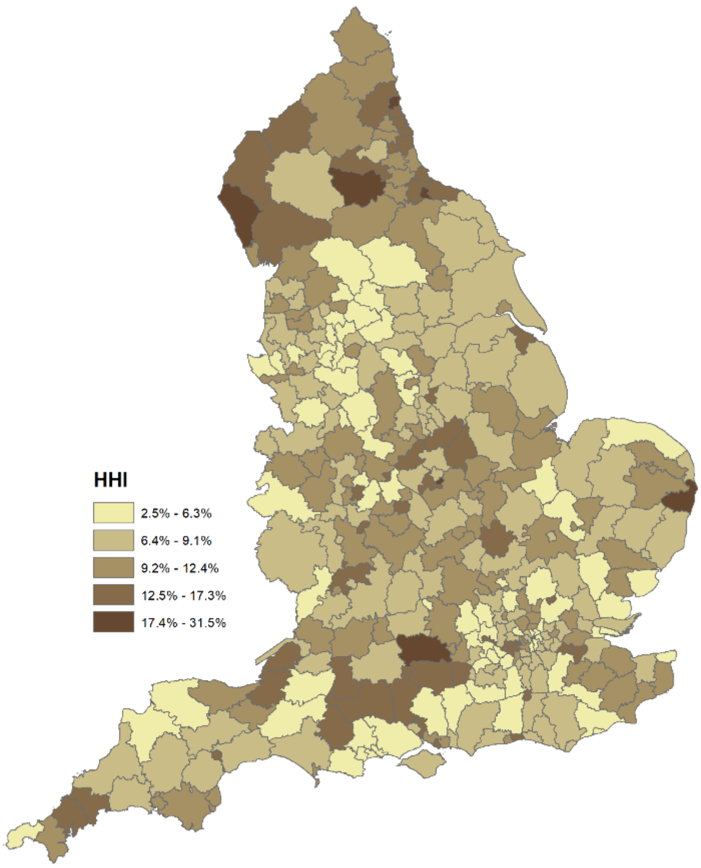


Fig. B7: Average Construction Duration (in Days) at the Dwelling Level

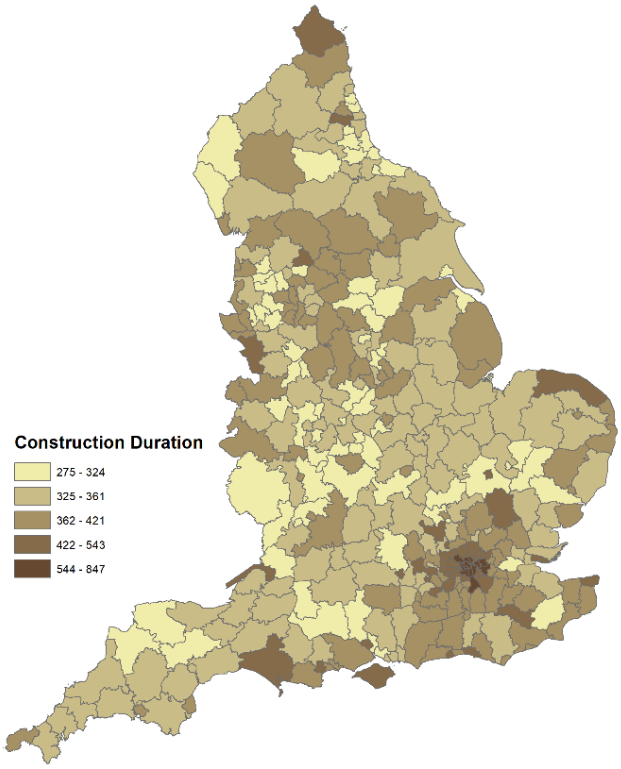
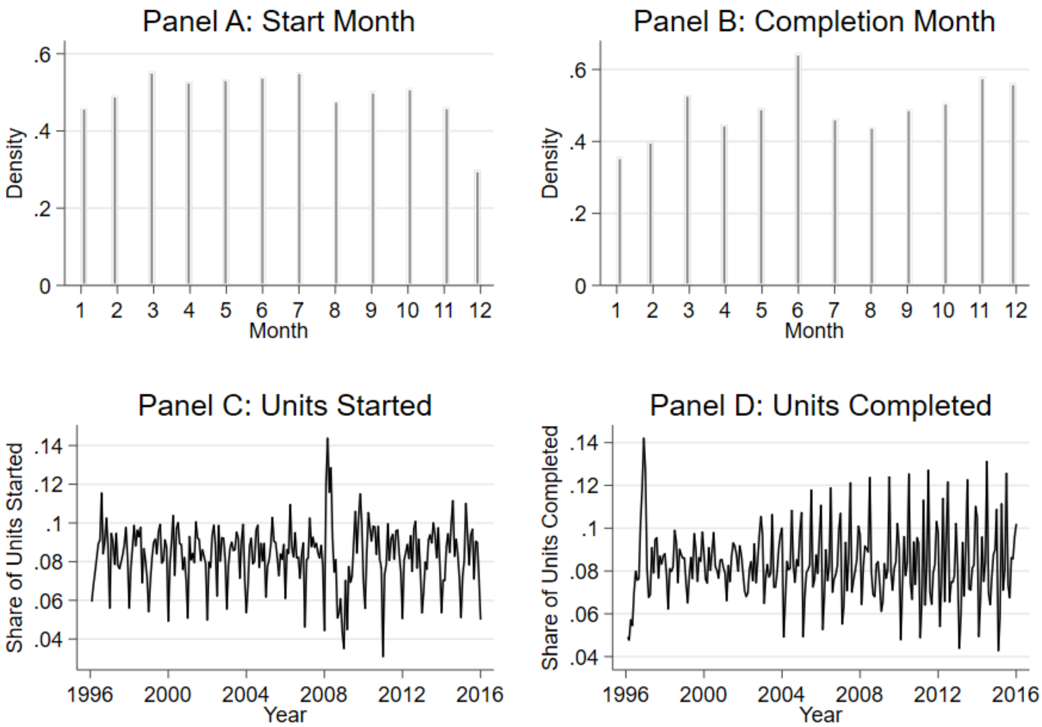


Fig. B8: Seasonality of Site Construction



Chapter 4

Supply Constraints and Housing Markets:

Evidence from a Spatially Matched Dataset

1. Introduction

A large and rapidly growing literature has investigated housing supply constraints and estimated their impacts on housing markets (Glaeser *et al.* 2005, Saiz 2010, Hilber and Vermeulen 2016, Brueckner *et al.* 2017, Brueckner and Singh 2020). Previous studies use different approaches to measure supply constraints and find that both regulatory and geographical constraints contribute to house price appreciation. Over the past decades, China has experienced a rapid process of urbanization and house prices have risen staggeringly in major cities. However, due to a lack of reliable measures, few studies have explored the role of supply-side factors in China's housing markets.

This paper first provides a comprehensive measure of regulatory restrictiveness in major Chinese cities. I apply and extend the approach proposed by Glaeser *et al.* (2005) to evaluate the 'regulatory tax' in 117 cities.[71] To do so, I combine 14,047 residential projects with land transaction records and construction cost information. The unique spatially matched dataset contains rich information on site size, land cost, construction cost, and housing price, which enables me to estimate a regulatory tax that can capture all the costs imposed from the government side. I then compute the average regulatory tax rate at the city level and find substantial spatial variations in regulatory restrictiveness across major Chinese cities. The comprehensive feature of this paper's measure makes it different from previous studies on housing supply constraints (Glaeser *et al.* 2005, Cheshire and Hilber 2008, Brueckner *et al.* 2017), which mainly focus on the costs imposed by the building height regulation.

I also utilize the regulatory tax estimate and a geographical scarcity variable, measured by the share of developed land relative to developable land in 2005, to explore the impact of supply-side constraints on housing markets. The main empirical challenge

[71] In this paper, a city refers to a 'prefecture', which is an administrative jurisdiction in China. Prefectures include superstar cities such as Beijing and lower-Tier cities such as Jiaxing.

faced by this paper is that housing supply constraints are likely to be correlated with local housing market conditions and other characteristics (e.g. corruption). To address the endogeneity concern, I apply an instrumental variable- and fixed effects-strategy. Exploiting the exogenous variations generated from local government tax enforcement, number of dialects, and the distance to historical amenities, I find that local housing demand shock, measured by the average real salary, will increase real housing prices, and this impact is stronger for places with more restrictive regulatory and geographical constraints. This paper's IV estimates suggest that a 1% increase in real salary will increase real house price by 0.17%. Besides, if an average city observes a one standard deviation increase in its regulatory tax rate, a 1% increase in real salary will increase real house price by 0.28% instead of 0.17%. In a similar vein, a one standard deviation increase in the geographical constraint will lead to a 0.24% instead of a 0.17% increase in real housing price. The main findings are robust when I apply a Bartik-type predicted local employment as the demand shifter and when I use two alternative estimates of the regulatory tax. I also find that local demand shock will stimulate more land supply and new construction, but less so for places with more restrictive supply constraints.

This paper's findings are important for three reasons: First, the rich information of the spatially matched dataset enables me to estimate a regulatory tax that captures all the costs coming from the government side. This measure is thus different from previous studies that mainly focus on the costs introduced by the building height regulation and provides additional insight into local regulatory restrictiveness.

Second, few studies have investigated housing supply constraints in China due to a lack of reliable measures. To the best of the author's knowledge, this paper is the first to provide a comprehensive estimate of regulatory restrictiveness that covers over 100 major cities in China.

Third, this paper proposes an alternative explanation for the staggeringly high housing price appreciation in major Chinese cities. The National Bureau of Statistics records that the average nominal housing price increases by 313% in 35 major cities between 2003 and 2018. Previous research suggests that house price growth in China is mainly driven by demand-side factors such as income growth, rapid urbanization (Garriga *et al.* 2020), and a deeper integration with the global market (Wang *et al.* 2011). However, few studies have explored the role of supply-side constraints in China's housing markets. This paper aims to fill this gap.

This paper relates to the literature discussing the measurement and consequences of supply constraints (Gyourko *et al*. 2008, Saiz 2010, Hilber and Vermeulen 2016, Brueckner *et al*. 2017). For instance, Gyourko *et al*. (2008) use the responses from a nationwide survey of residential land use regulation in over 2,600 communities across the US to develop a series of indexes that capture the stringency of local regulatory environments. Brueckner *et al*. (2017) develop a new approach for measuring the stringency of a major form of land use regulation, building height restrictions, and applies it to a dataset of land-lease transactions from China. Glaeser *et al.* (2005) propose that the gap between observed housing market price and the marginal cost of building up an additional floor could be interpreted as a 'regulatory tax' imposed by local administrations. Cheshire and Hilber (2008) and Cheung *et al.* (2009) follow this method to evaluate regulatory tax in the UK and in the US, respectively. As previous studies mainly focus on the regulatory restrictiveness imposed by the building height regulation, this paper contributes to the literature by estimating a regulatory tax that captures all the costs imposed from the government side.

This paper also ties into the literature discussing real estate markets in China (Wu *et al.* 2012, Wu *et al.* 2016, Fang *et al.* 2016, Glaeser *et al.* 2017, Liu and Xiong 2018). Previous studies propose that demand-side factors such as income growth and rapid urbanization are driving up housing prices. This paper contributes to the discussion by providing an alternative perspective from the supply side.

The rest of this paper is structured as follows. Section 2 describes the details of housing markets and real estate development in China. Section 3 presents the data sources, descriptive statistics, and how I estimate the regulatory tax rate. Section 4 discusses the identification strategy and presents main empirical results. Section 5 concludes.

2. Institutional Background

2.1. Housing Markets in China

Chinese cities have been urbanizing rapidly since the major economic reform in 1978. According to the National Bureau of Statistics, the share of the urbanized population in China rises from 25.8% in 1990 to 59.6% in 2018. Followed by a national housing market reform in 1998, real estate development has become a booming industry and housing prices have raised staggeringly in major cities. As shown in Figure 1, the average nominal housing price in Tier-1 cities has raised from 6,224 RMB per square meter to 35,732 RMB per square meter between 2005 and 2018, with an annual growth rate of 12.7%. Even for a group of Tier-2 cities, the average nominal housing price has increased from 2,840 RMB per square meter to 11,656 RMB per square meter during the same period, with an annual growth rate of 9%.[72]

There is a heated debate in China about whether the house price appreciation is driven by strong economic growth or irrational exuberance. As shown in Panel A of Figure 2, while the ratio of average salary relative to the average housing price has been increasing in Tier-2 cities and lower-Tier cities during 2005 and 2018, it has been declining in Tier-1 cities during the same period, suggesting that housing becomes less and less affordable in superstar cities. Meanwhile, Panel B of Figure 2 shows that the ratio of construction cost relative to house price has been decreasing in major cities between 2005 and 2018, which indicates that the house price appreciation is not mainly driven by the increase in construction cost in these cities.

An alternative explanation for high housing prices in major Chinese cities is the lack of supply. As shown in Figure 3, the average land supply per person in Tier-1 cities has become lower than other major cities since 2008 even though demand is much stronger in these superstar cities. To investigate the role of supply-side factors in China's housing markets, it is important to understand the land supply system and the process of real estate development, which will be discussed as below.

2.2. Land Supply and Real Estate Development in China

Since 1988, the city land bureau has gotten the authority to allocate the use rights (leaseholds) of vacant urban land plots. After 2004, all urban land leases for private development need to be allocated through public auctions in order to control for the corruptions occurred during the negotiations between local governments and developers. Developers bid for land plots in the auction markets and then build out residential projects upon the land parcels. A typical residential project in China

[72] Tier-1 cities refer to Beijing, Shanghai, Guangzhou, and Shenzhen. There are different classifications of Tier-2 cities (or 'New Tier-1 cities' in some cases), but in general these classifications include similar cities. In this paper, Tier-2 cities refer to all sub-provincial ('fushengji') cities and two municipalities ('zhixiashi').

contains hundreds to thousands of dwellings and can be constructed by either one or several developers. After acquiring land plots, developers need to make construction plans and apply for different approvals from the local governments. Once a construction permit is granted, developers can start to build out sites and will try to meet governments' requirements to list properties. A national policy regulates that developers are permitted to list properties only if they achieve at least 25% of their investment and have a detailed construction and completion plan. After getting the pre-sale permit, developers can sell properties, continue building out sites till the completion, and transfer properties to homebuyers.

During the real estate development process, developers are faced with different types of costs coming from the government side. First, there is a cost related to the inefficiency of local administration. Applying for planning and construction approvals requires a significant amount of time and efforts. For instance, according to a proposal to the People's Congress of Hainan Province (Hainan Provincial Government, 2014), there are over 30 approvals that need to be applied for during the development process, which takes 272 approval days. Some local governments have tried to improve the administration efficiency of granting approvals. For instance, Hainan province has optimized the approval process and reduced the total approval time from 799 days to 37 days (Sichuan Provincial Government, 2014). However, inefficient administration remains as a major cause for construction delay in some cities and introduces additional costs for developers. Second, different types of taxes and administrative fees are required to be paid during the development process. According to an interview with a Congress representative, there are as many as 12 taxes and 50 fees related to the real estate industry, and these taxes and fees all together account for between 30% and 40% of real estate prices based on a sample survey (Tencent News, 2010). Third, some cities in China introduce restrictive land use regulations such as low floor area ratio (FAR) limits. Developers need to either comply with these regulations or negotiating with local governments, which further increases development costs. This paper proposes a regulatory tax measure that captures all these costs coming from the government side. The details of the data and the method are discussed in section 3.

3. Data and Descriptive Statistics

3.1. Data and Spatially Matched Residential Projects

This paper uses two main datasets for the empirical analysis. The first one comes from Fang.com, one of the leading real estate agencies in China. It provides information about over 50,000 residential projects in 150 major cities. For each residential project, the dataset records average house price per square meter in 2018, the land area of the project, the total construction area, floor area ratio, developer, address, and the time when the developer first lists properties. Each project is geocoded using its address information. Another main dataset used in this paper covers over 200,000 residential land transactions in 287 Chinese cities between 2005 and 2018. The data source is the official website of China land market. The land transaction dataset records detailed information at the plot level including land transaction price, address, the date of transaction, the upper limit and the lower limit of floor area ratio, land area, the type of auction, and the land bidder. Each land plot is geocoded using its address information.

I follow two steps to spatially match these two datasets and estimate the land cost for each residential project: First, as shown in Figure 4, I draw a buffer with a 3 km radius for each residential project and select all the land plots that are within this buffer and released in the same year when the first dwelling of the residential project is listed. Second, I compute the average land price per square meter for all the selected land plots within the 3-km-radius buffer. These two steps allow me to have an accurate estimate of the land cost for each residential project based on the land transaction records nearby.[73] I then follow the method that will be discussed in section 3.2 to estimate regulatory tax rate at the project level. I drop the top 1% and the bottom 1% of observations in terms of regulatory tax rate to mitigate the potential bias caused by extreme values. In the baseline estimation sample, I also drop projects with negative regulatory tax rates and cities with fewer than 10 matched pairs.

In addition, this paper applies a stricter one-to-one matching method by selecting residential projects and land plots that are geographically close to each other, have a similar land area, and share the same name between the land bidder and the project developer. In this scenario, each residential project is accurately matched with its corresponding land plot. This one-to-one matching dataset is not used in the main

[73] The average land price per square meter equals the land transaction price divided by the 'jirong' area, which is defined as the land area multiplied by the upper FAR limit. I use the 'jirong' area instead of the land area to compute land price per square meter so that I can take into account the effect of FAR design on land value.

analysis because only 2,476 projects can meet the strict criteria. However, sections 3.3 and 4.4 will show that both matching methods have similar estimates of regulatory tax rates, and the empirical results are robust regardless of the matching methods applied.

This paper also collects city-level characteristics from different sources including China City Statistical Yearbook, China Financial Statistics of Cities and Counties, China Real Estate Statistical Yearbook, China Financial Statistical Yearbook, China Historical Geographic Information System, and prefecture-level statistical yearbooks. I also collect global land cover maps from the European Space Agency to identify local built-up areas. The financial information of 60 major developers is acquired from Orbis to estimate developer profit level.

Basic summary statistics computed for the matched dataset of land plots and residential projects are detailed in Panel A of Table 1. There are in total 14,047 matched residential projects between 2005 and 2018. The average house price is 11,585 RMB (around 1,274 GBP[74]) per square meter, the average land cost is 2,612 RMB per square meter, and the average construction cost is 2,831 RMB per square meter. These residential projects are usually large in size and contain 1,242 dwellings per site on average. The average regulatory tax rate at the project level is 1.7. Panel B of Table 1 shows the descriptive statistics for a panel of 105 cities. Both the nominal house price and nominal salary are deflated by the national retail price index in order to obtain a real house price and salary at the 2005 level. The average house price is 4,678 RMB per square meter and the average annual salary is 36,824 RMB. The average regulatory tax rate at the city level is 1.6. Panel C reports the descriptive statistics for the city-level instruments, including fiscal revenue relative to GDP in 1997, number of dialects within a city, and the distance to major lakes and rivers in 1820. I will discuss my identification strategy in detail in section 4.

3.2. Estimate Regulatory Tax Rate in China

As a measure of regulatory restrictiveness, the concept ‘regulatory tax’ is first proposed by Glaeser *et al.* (2005). The key idea is quite intuitive: in a housing market with perfect competition among developers and free entry and exit, housing price per square meter will equal to the construction cost per square meter. Since the marginal construction cost rises with building height, in the absence of height limits, buildings

[74] Based on the currency exchange rate in September 2021.

should rise to a point where the marginal construction cost of adding an extra floor equals the market housing price. Any gap between the market housing price and the marginal construction cost can be interpreted as a 'regulatory tax', which represents the costs for developers to negotiate with local governments and get the planning approval to add an additional floor.

I follow and extend this approach to estimate regulatory tax in the context of China's housing markets. Table A1 provides a breakdown of different types of development costs from an anonymous developer in China. As the table suggests, the costs of real estate development in China can be generally categorized into 3 items: land cost (the price that developers pay to bid for the land plot), construction cost (the cost of labour and raw materials to build out sites), and regulatory tax (administrative fees and taxes during the development process, developers' commitments to provide public facilities, the costs of inefficient administration, and the costs for developers to comply with restrictive land use regulations). Under a market with perfect competition among developers, the total sales of a residential project should equal to the total costs. Regulatory tax can thus be computed by:

$$\text{Regulatory Tax} = \text{Total Sales} - \text{Land Cost} - \text{Construction Cost} \tag{1}$$

To reflect the fact that developers normally expect some profits, I estimate a profit margin by using a sample of 60 major developers in China and include a 'profit' component in equation (1). Following Cheshire and Hilber (2008), I divide the regulatory tax by construction cost to normalize this measure. Regulatory tax rate can thus be computed by:

$$\text{Regulatory Tax Rate} = \frac{\text{Total Sales} - \text{Profit} - \text{Land Cost} - \text{Construtcion Cost}}{\text{Construtcion Cost}} \tag{2}$$

The spatially matched dataset provides sufficient information for me to estimate equation (2) at the project level. I compute the total sales by multiplying the average house price per square meter with the total sellable area of the project.[75] The estimated profit margin based on 60 major developers in China suggests that on average,

[75] The dataset provides the average house price per square meter of each residential project in 2018. I adjust the house price to the year when the residential project starts to list properties using city-level price index in order to stay consistent with the year to estimate land cost and construction cost. Total sellable area is defined as the total land area of a residential project multiplied by its floor area ratio. This is a more realistic estimate of the total sellable area compared with the total construction area, which includes unsellable areas such as driveways, heating boiler rooms, the basement used for civil air defense, etc.

developer's net profit equals to 11% of total sales. Therefore, I estimate the profit of each project as 11% of the total sales. I also estimate the land cost by multiplying the average land price per square meter with the total sellable area. Finally, I compute the construction cost of each project by multiplying the total construction area with the average construction cost per square meter. The data for construction cost per square meter is time-varying at the city level. Every year, developers are required to report their total construction area and total construction cost to local governments. The construction costs data is computed based on the information provided by developers, and this measure includes the costs of both materials and labours.[76]

The regulatory tax rate computed from equation (2) captures a series of taxes and administrative fees that occur during real estate development. In addition, if a local government introduces restrictive land use regulations such as low FAR limits, the costs for developers to comply with or negotiate to change these regulations will be captured into the regulatory tax rate measure. The construction delay caused by inefficient administration will be included in the measure as well. The comprehensive feature of this paper's regulatory tax estimate makes it different from the measures in previous studies (e.g. Glaeser *et al*. 2005, Cheshire and Hilber 2008, Bruekner *et al*. 2017), which mainly focus on the costs introduced by the building height regulation. I then compute the average regulatory tax rate in 117 cities to get a measure of local regulatory restrictiveness. To the best of the author's knowledge, this paper is the first to provide a regulatory constraint measure that covers over 100 major cities in China.

This paper also conducts two alternative estimates of regulatory tax rate: The first one is computed based on the stricter one-to-one matching sample. The second one is estimated by following the literature and assuming perfect competition among developers. In this case, developers will make zero profit, and the regulatory tax rate is estimated as below:

$$\text{Regulatory Tax Rate (Alternative)} = \frac{\text{Total Sales} - \text{Profit} - \text{Land Cost} - \text{Construtcion Cost}}{\text{Construtcion Cost}} \quad (3)$$

3.3. Descriptive Statistics of Regulatory Tax Rate

[76] In my dataset, occasionally the construction cost data is not available for some cities in some years. In this case, I impute the missing value by taking the average of the construction cost per square meter in this city between 2005 and 2018.

Following the estimation strategy as discussed in section 3.2, this paper first computes the regulatory tax rate at the project level. The distribution of the estimated regulatory tax rate is presented in Figure 5. There are in total 14,047 projects and most of them have a regulatory tax rate below 5. Figure 5 also shows that the residential projects in my estimation sample usually contain many dwellings. By simply aggregating the 10,999 projects with information about the number of dwellings, I find that the matched dataset covers in total 13,657,338 dwellings, which provides a good coverage of residential developments in major Chinese cities.

I then compute the average regulatory tax rate at the city level to measure local regulatory restrictiveness and the results are shown in Table 2. Three Tier-1 cities, Beijing, Shanghai, and Guangzhou, have the top 3 highest regulatory tax rates among 117 cities. This is in line with expectation because these cities tend to design more restrictive land use regulations and thus introduce additional costs for real estate development. For instance, Beijing has a strict building height limit in the central area to protect historical architectures and cultural heritages. Besides, this paper's estimated regulatory tax rate also captures taxes that are computed based on local land value and housing price, which tend to be high in superstar cities. The regulatory tax estimates in Beijing and Shanghai are also comparable with the estimates in major European cities.[77]

However, it is not always the case that cities with high housing prices also have high regulatory taxes. For instance, Shenzhen, another Tier-1 city in China with the highest average housing price among all Chinese cities in 2018, has a regulatory tax rate of 1.8, which is just above the sample mean value of 1.6 and is ranked as the 38th of all 117 cities. The relatively low regulatory tax rate in Shenzhen, despite of high housing price there, is also reasonable, as Shenzhen is a fast-growing city and is famous for its efficient administration and the so-called 'Shenzhen speed' for new development.

The rich information of the matched dataset also allows me to disentangle the relative importance of different components that contribute to house price (regulatory tax, land cost, construction cost, and developer profit). Figure 6 presents the decomposition of

[77] Cheshire and Hilber (2008) estimate regulatory tax rate in major European cities. For instance, in 2005, the regulatory tax rates in City of London, Stockholm, Frankfurt, and Barcelona are 4.31, 3.3, 3.31, and 3.16 respectively. These estimates are comparable with the regulatory tax rates in Beijing (5.42), Guangzhou (3.71), and Shanghai (3.32).

house prices in four Tier-1 cities and Jiaxing, which is a medium-size city close to Shanghai and is reckoned as an example of cities with relaxed supply constraints. The decomposition in Figure 6 shows that regulatory tax accounts for the largest proportion of house price in Tier-1 cities among all the components. The ratio of regulatory tax relative to house price is 50.7% in Beijing, 50.9% in Shanghai, 49.6% in Guangzhou, and 39.3% in Shenzhen, respectively. Conversely, construction cost and land cost contribute more to the house price in Jiaxing, where land use regulations are less restrictive compared with Beijing and Shanghai. Figure 7 then presents the average house price decomposition in Tier-1 cities, Tier-2 cities, and other cities in China. We can observe that regulatory tax accounts for a larger proportion of house price in Tier-1 cities compared with lower-Tier cities. At the nation level, on average, regulatory tax, land cost, and construction cost account for 38.6%, 19.7%, and 30.7% of house price, respectively.

One might argue that developer profit margin can be higher in superstar cities, leading to an overestimate of the regulatory tax rates there. However, data from the National Bureau of Statistics suggests that in fact, there is a higher degree of developer competitiveness in Tier-1 cities compared with lower-Tier cities. For instance, the ratio of the average number of developers between 2005 and 2018, this paper's main estimation period, relative to the population (1,000 people) in 2005 is higher in Beijing (0.24%) and Shanghai (0.23%) compared with this ratio in Tier-2 cities Chongqing (0.07%) and Tianjin (0.13%). If there is a higher degree of competitiveness in Tier-1 cities, developers are expected to make fewer profits there, and this paper's measure might even underestimate the regulatory restrictiveness in these superstar cities.

Although no rigorous academic research has estimated regulatory tax rate in China before, China Real Estate Chamber of Commerce has conducted a survey on 81 real estate projects in 9 major cities in 2008 with the aim to understand why housing prices are so high in some cities (Sohu Finance, 2009). The survey suggests that on average, land cost, taxes, and fees account for 49.42% of total sales. This ratio is higher in Shanghai (64.5%) and Beijing (48.3%). Their findings are in line with this paper's estimates and suggest that high regulatory cost is one of the main reasons for high housing prices in major cities.

Figure 8 presents regulatory tax rate at the city level and shows substantial spatial variations in regulatory restrictiveness across 117 Chinese cities. Cities along the

South-eastern coast such as Shanghai and some core cities in the North such as Beijing have higher regulatory tax rates compared with other cities. Regarding the accuracy of this paper's measure, Panel A of Figure 9 presents the correlation between the regulatory tax rate based on the estimated land value sample and the regulatory tax rate based on the one-to-one matching sample. The correlation coefficient is high and positive. Besides, Panel B of Figure 9 shows that there is a positive correlation between population and local regulatory tax rate, suggesting that larger cities tend to be more restrictively regulated.

4. Empirical Analysis

In this section, I explore the impacts of supply-side constraints on housing markets in major Chinese cities between 2005 and 2018. I use the period between 2005 and 2018 because it is the same period as the spatially matched dataset covers.

Figure 10 presents a simple theoretical intuition to guide the empirical analysis. Panel A shows a housing market with restrictive supply conditions. In this market, a strong local demand shock will not stimulate much construction, and housing price will increase significantly. Conversely, Panel B presents a housing market with lax supply constraints and shows that a strong demand shock will lead to a significant increase in housing construction and housing price will stay relatively stable.

This paper tests the following proposition in the empirical analysis:

Proposition – Positive local demand shocks increase housing prices and construction. The demand effect for housing prices is stronger in cities with more restrictive supply constraints, and the demand effect for construction is stronger in cities with less restrictive supply constraints.

4.1. Empirical Specification and Identification Strategy

To test for the main proposition, I apply two variables that can measure local supply conditions: Frist, the regulatory tax rate as discussed in section 3.3 provides a comprehensive measure of local regulatory restrictiveness at the city level. Second, I use the share of developable land that is already developed in 2005, the beginning year of my sample period, as a measure of local geographical constraint for cities to expand horizontally.[78] This paper then estimates the following specification using OLS:

$$Ln(House\ Price_{ct}) = \beta_1 Ln(Salary)_{ct} + \beta_2 Ln(Salary)_{ct} \times RTR_c + \beta_3 Ln(Salary)_{ct} \times Geography_c + \phi_c + \delta_t + \varepsilon_{ct} \quad (4)$$

where c indexes each city and t indexes year. A vector of city fixed effects is represented by ϕ_c. δ_t is a set of time dummies (year fixed effects). The dependent variable represents the natural logarithm of real housing price in city c and year t. The variable $Ln(Salary_{ct})$ represents the natural logarithm of average real salary in city c and time t. Both the average housing price and the average salary measures are adjusted to the 2005 level using the national retail price index. RTR_c represents the regulatory tax rate in city c, and $Geography_c$ denotes the share of developable land that is developed in city c in 2005. The parameters of interest are β_2 and β_3, measuring the impact of supply constraints on house price-salary elasticity. Both supply constraints measures are standardised so that I could easily interpret the estimated coefficients. I subtract the sample mean of each measure from the measure itself and divide this difference by the standard deviation of the measure. This transformation allows me to interpret the estimated coefficient as an increase in the price-salary elasticity due to a one standard deviation increase or decrease in the supply constraint measure.

One important caveat with the OLS estimates of equation (4) is that the key explanatory variables are likely endogenously determined. First, there is a concern of reverse causality. Regulatory tax rate captures taxes that are correlated with local house prices. If regulatory tax rate is positively correlated with housing price, the OLS estimate will underestimate the impact of regulatory constraint on house price-salary elasticity. Second, unobserved spatial and administrate features such as local amenities, corruption, and government quality might influence the share of developed land, regulatory tax rate, and housing prices simultaneously. To addresses these endogeneity issues, this paper first control for city fixed effects and year fixed effects to mitigate the concerns of unobserved time-invariant local factors and the macro trends. However, the concerns of reverse causality and unobserved time-varying confounding features are still not fully addressed.

[78] Figure B1 presents the share of developed land for the 105 cities used in my estimation.

This paper then proposes three sets of instruments for the supply constraint measures to address these endogeneity issues. The first instrument is the number of dialects within a city. The identifying assumption is that more dialects within a city will reduce the efficiency of communication between local governments and developers and thus increase the administrative costs and regulatory tax. However, there is no direct correlation between the number of dialects and the current house prices, as dialects are usually formed by historical and cultural features.[79] The second instrument for the regulatory tax rate is the level of local tax enforcement in 1997. Real estate industry in China starts to be marketized after a major housing reform in 1998, and before 1998, local fiscal revenue was lightly relied on housing markets. I use the local fiscal revenue relative to local GDP in 1997 as a proxy for local tax enforcement before the development of real estate industry in China. The identifying assumption is that local governments' ability to collect taxes and administrative fees is positively correlated with regulatory tax rate, but it doesn't have a direct impact on housing prices today because real estate industry had a negligible influence on local fiscal revenue and GDP before 1998.

I also propose an instrument for the share of developable land that is developed in 2005. As shown in Figure B3, I compute the distance between a city's geometric centroid and the major lakes and rivers in 1840 and apply this distance as an instrument for the geographical constraint measure. The identifying assumption is that places close to the major rivers and lakes in 1840 tend to be developed earlier and thus have a higher share of land developed in 2005. However, the distance to the historical amenities should not have a direct and systematic influence on house price today, after I control for city fixed effects in the specification.

4.2. Main Results

Table 3 summarizes the main findings for estimating equation (4). Column (1) reports the results for the naïve OLS specification controlling for both year and city fixed effects. The coefficient on the house price-salary elasticity is highly statistically significant and positive, suggesting that in a city with average supply constraints, a 1% increase in local real salary increases local real housing price by 0.18%. The coefficients on the salary-interactions with the regulatory tax rate and the share

[79] Figure B2 presents the number of dialects for 105 cities that are included in the baseline estimation.

developed-measure are also positive and statistically significant, implying that if an average city observes a one standard deviation increase in its regulatory tax rate, a 1% increase in local salary will increase house price by 0.21% instead of 0.18%. In a similar vein, a one standard deviation increase in the share developed land leads to a 0.26% instead of a 0.18% increase in housing price.

The OLS specification ignores endogeneity concerns about local regulatory restrictiveness and the share of developed land. In column (2) of Table 3, I estimate the same regression by Two-Stage Least Squares (2SLS), instrumenting the regulatory tax rate-salary and the share developed-salary interactions. Consistent with the OLS estimates, the coefficients on the real salary, the regulatory tax rate-salary, and the share developed-salary interactions are all positive and statistically significant. While the IV estimated coefficient of the share developed-salary interaction is similar to the OLS estimate, column (2) reports a larger impact of regulatory tax rate on house price-salary elasticity compared with the OLS result, suggesting that a 1% increase in local salary will increase house price by 0.28% instead of 0.17%. This is in line with the expectation that reverse causality will cause an underestimate of the variable of interest and the instrumental variable strategy can mitigate this endogeneity concern.

Regarding the validity of the instruments, the Kleibergen-Paap F statistic indicates that weak identification is not a problem. I also report the first-stage regression results in Panel B of Table 3. Column (3) of Table 3 suggests that both local tax enforcement and the number of dialects are positively and significantly correlated with regulatory tax rate. In addition, column (4) suggests that the distance to major rivers and lakes in 1840 has a negative and statistically significant correlation with the share of developed land. All first-stage results in Table 3 show that the instruments are relevant to the endogenous variables in expected ways.

4.3. Quantitative Analysis

This section presents a counterfactual analysis based on the IV results reported in Table 3. The TSLS specification yields a prediction of real house price conditional on the real salary, regulatory constraint, geographical constraint, as well as city and year fixed effects. Based on the estimated coefficients, I first obtain counterfactual scenarios by predicting local real housing price with supply constraints set to zero sequentially. I then also remove the independent effect of real salary, in order to identify the counterfactual housing price holding constant all relevant local demand and supply measures. This exercise allows me to understand the quantitative importance of the variables of interest.

Removing all supply constraints in housing markets are of course unrealistic scenarios in practice. Therefore, I explore an alternative exercise, where I remove one standard deviation of regulatory constraint and geographical constraint sequentially. I then also remove the independent effect of the real salary to obtain the predicted housing price for a city absent of fluctuations in local housing demand. I first conduct this exercise for each city separately and then take the average of the predicted housing prices over all cities to derive a counterfactual scenario for the 'average' Chinese city. To explore the relative importance of geographical and regulatory constraints, I also conduct two exercises by separately removing these two supply constraints, and by separately lowering them by one standard deviation.

The results of these quantitative exercises are summarized in Table 4. The corresponding Figures 11 and 12 illustrate the predicted housing prices between 2005 and 2018 for the 'average' Chinese city under two scenarios: variables of interest set to zero and reduced by one standard deviation. Figure 13 illustrates the scenarios for a few distinctive cities with tight or comparably relaxed planning constraints: Beijing and Shanghai are cities with more restrictive land use regulations. Shenzhen is a Tier-1 city with high housing price, geographical scarcity, but relatively lax regulatory environment. Jiaxing is a medium-size city with few regulatory and geographical constraints.

The counterfactual exercises suggest a substantial impact of supply constraints on house prices in China. Panel A of Table 4 suggests that, based on the baseline estimates, in 2018 the house price in the 'average' city in China (with average housing demand shocks) would be 20.6 percent lower if all regulatory constraints were completely relaxed. Panel B then shows that reducing the restrictiveness by one standard deviation were to lead to a 10.3 percent reduction in housing prices. If I completely remove both supply-side constraints and the independent effect of real salary, the housing price in the 'average' Chinese city in 2018 would be 35.8 percent lower. Regarding the relative importance of geographical and regulatory constraints, both panels C and D suggest that reducing the regulatory constraint would have a larger quantitative impact on housing prices compared with reducing the geographical constraint in the ‘average’ Chinese city.

As Figure 13 illustrates, the impacts of regulatory constraints and physical (scarcity related) constraints vary significantly across locations. Physical constraints matter most in the densely developed city, Shenzhen, while regulatory constraints are most important in Beijing and Shanghai. In Jiaxing, both regulatory and geographical constraints have relatively small impacts on housing price, and the independent effect of real salary is more important for house price appreciation.

4.4. Additional Results and Robustness Checks

4.4.1. Supply Constraints, Land Release, and Housing Construction

In this sub-section, I study the impacts of supply constraints on housing construction and the release of land plots by changing the dependent variable in equation (4) and re-estimating the specification.

Columns (1) and (2) of Table A2 report the OLS and IV estimates for the impact of supply constraints on the release of land parcels by local governments. Both columns show that real salary has a positive and statistically significant impact on local land supply. The preferred IV estimate in column (2) suggests that the share of developed land has a negative and significant impact on land supply-salary elasticity. This is in line with expectation as local governments will release more land plots when they observe high housing demand, but less so for cities with scarcity in developable land. Column (2) also reports an insignificant effect of the regulatory tax rate on the land supply-salary elasticity. This is reasonable as regulatory tax will only matter when the real estate development has started, and the land plot has already been released. Columns (3) and (4) of Table A2 report the OLS and IV estimates for the impact of supply constraints on housing construction. Both columns show that real salary has a positive and statistically significant impact on the total area of housing under construction. The preferred IV estimate in column (4) suggest that regulatory tax rate has a negative and significant impact on housing construction-salary elasticity. Columns (2) and (4) together suggest that both local governments and developers will respond to strong demand shocks by releasing more land plots and by building more housing, respectively, but this demand effect is weaker for cities with more restrictive supply constraints.

4.4.2. Labour Demand Shock as Demand Shifter

In the baseline estimation, I use city-level real salary as the local housing demand shifter. However, this measure might be endogenously determined, as housing supply conditions might directly affect local labour market and thus local salary (Saks 2008). To address this endogeneity concern, this paper follows the method proposed by Bartik (1991) to construct a shift-share type predicted local employment measure that could be used as local housing demand shifter. I exploit the city-level variation in industry composition of employment in 2003, the earliest year with available data, and the changes in employment by industry at the national level between 2005 and 2018 to construct this labour demand shock variable.

Figure B4 presents the national trend of employment by industry series, where the 2003 level has been set to 1. There are in total 19 industries as classified by the National Bureau of Statistics in China, and Figure B4 highlights a significant rise of employment in the information technology sector as well as in the finance sector, and a decline of the mining and agriculture employment between 2003 and 2018. For each city and industry, employment was multiplied with the corresponding national trend and the result was aggregated over industries. This yields the predicted employment in each city that would have resulted given its industry composition in 2003, if employment in each industry follows the national trend. This paper argues that this labour demand shock variable is exogenous to local housing demand and supply conditions, as local labour supply shocks have a negligible impact on industry employment at the national level.

This paper then applies the natural logarithm of the labour demand shock variable (LLDS) as the demand shifter and re-estimates equation (4). The results are reported in Table A3. Columns 1 and 2 present the OLS and IV estimates respectively, and both columns report positive and statistically significant coefficients for the regulatory tax rate-LLDS and the share developed-LLDS interactions, which are in line with the baseline estimation results. This exercise suggests that this paper's main findings are robust after I mitigate the potential endogeneity concern of the local demand shifter.

4.4.3. Alternative Estimates of Regulatory Tax Rate

The regulatory tax rate used in the main analysis is computed based on the estimated land price within the 3 km radius buffer and the assumption that developers will make ‘normal’ profits. In this sub-section, I change these two assumptions separately and estimate alternative regulatory tax rates to test for the robustness of the main results.

I first use the one-to-one matching sample to re-estimate regulatory tax rate at the city level. As discussed in section 3, the one-to-one matching sample contains residential projects with their corresponding land plots. This sample thus provides a more accurate estimate of the land cost. The drawback of using this sample is that the city-level regulatory tax rate is estimated based on fewer observations due to the stricter selection criteria. The one-to-one matching sample provides an alternative estimate of regulatory tax rate for 60 cities and I re-estimate equation (4) using this alternative measure. The results are reported in columns (1) and (2) of Table A4. In line with the baseline findings, all the estimated coefficients have the expected signs and are statistically significant.

I then follow the literature and assume perfect competition among developers when estimating regulatory tax rate. Under this scenario, developers will make zero profit, and the regulatory tax rate is estimated based on equation (3). I then apply this alternative measure of regulatory tax rate to re-estimate equation (4). The results are reported in columns (3) and (4) of Table A4 and are consistent with the main findings.

4.4.4. Remove Tier-1 Cities and Municipalities

In the end, this paper conducts a robustness check by removing four Tier-1 cities (Beijing, Shanghai, Guangzhou, Shenzhen) and two municipalities (Tianjin, Chongqing) in the baseline sample and re-estimating equation (4). These six cities are more economically advanced in China and might capture unobserved spatial or administrative features that can bias the estimates. The results using this restricted sample are reported in Table A5. In line with the main findings, the coefficients of both interaction terms are positive and statistically significant, suggesting that the impacts of supply constraints are robust after I mitigate the concern of unobserved features in superstar cities.

5. Conclusion

This paper first estimates regulatory tax rate in China by using a unique spatially matched dataset and finds substantial variations in regulatory restrictiveness across major cities. I then apply the regulatory tax rate measure to study the impact of supply constraints on housing markets in China. I find that local demand shock will increase house price, and this demand effect is more pronounced for places with more restrictive regulatory and geographical constraints. Compared with previous studies with a focus on demand-side factors, this paper provides additional insight into understanding the staggeringly high house prices in major Chinese cities. This research topic is becoming more and more important for the global economy, as real estate value in China accounts for the largest proportion of the global property value.[80]

Figure B5 measures housing affordability at the city level in China in 2018 by using the ratio of average salary relative to housing price per square meter. The figure shows that it will take at least 8 years to buy a 100 m^2 property in Tier-1 cities and cities along the south-eastern coast for a homebuyer with an average salary. Based on international standards, housings in many Chinese cities can be classified as 'severely unaffordable'. [81] To stimulate more housing construction and to resolve the affordability issues in these cities, local governments can reduce regulatory tax by improving administration efficiency and lowering unnecessary taxes and fees related to real estate development. However, it is also worth noting that real estate industry is crucial for local economic development in China during the past decades and many local governments' budgets are heavily relied on fiscal revenues generated from the land and real estate markets. How to balance the pros and cons of high regulatory tax rates so that cities can develop in a sustainable way is an important question that can be explored in the future.

[80] As of 2016, China's real estate value is the largest in the world, accounting for 21% of global property value (Tostevin, n.d.).

[81] According to the 15th Annual Demographia International Housing Affordability Survey, cities with median multiples higher than 5 are classified as 'severely unaffordable'. Median multiple is defined as median house price divided by median household income.

Tables

Table 1: Descriptive Statistics

	Obs.	Mean	SD	Max	Min
Panel A: Matched project characteristics					
Average house price (RMB/m^2)	14047	11585.3	9007.8	194258	1705.9
Average land price (RMB/m^2)	14047	2612.1	3320.7	54040	101.2
Average construction cost (RMB/m^2)	14047	2830.6	898	9289	839
Regulatory tax rate	14047	1.7	2.2	18.7	0
Land area (m^2)	14047	88379.7	115324.1	999900	1044
Construction area (m^2)	14047	220958.1	260496.2	3500000	2475
FAR	14047	2.9	1.5	10	0.1
Number of dwellings	10999	1241.7	1417	30000	1
Panel B: City-level characteristics					
House price in 2005 level (RMB/m^2)	1444	4677.8	3596.6	43080.1	869.6
Average salary in 2005 level (RMB)	1444	36823.5	15716.6	116434.7	8051.3
Regulatory tax rate	1444	1.6	0.7	5.4	0.5
% Developed land in 2005 relative to developable land	1444	0.1	0.1	0.4	0.003
Predicted local employment	1456	881706	1411532	15224861	47621
Land supply (km^2)	1332	21.8	20.4	182.3	0.8
Housing construction area (km^2)	916	28.6	25	202.9	1
Panel C: City-level characteristics, instruments					
Local tax enforcement (fiscal revenue/GDP in 1997)	105	0.04	0.02	0.13	0.007
Number of dialects	105	2	1.1	6	1
Distance to major lakes and rivers in 1820 (km)	105	64.4	61.8	271.3	0.3

Table 2: Regulatory Tax Rate at the City Level

Ranking, city, and regulatory tax rate		
1. Beijing 5.419	41. Xi'an 1.796	81. Dalian 1.202
2. Guangzhou 3.71	42. Jinhua 1.795	82. Ningbo 1.196
3. Shanghai 3.315	43. Jinan 1.761	83. Changzhou 1.192
4. Fuzhou 3.243	44. Guiyang 1.758	84. Zhuhai 1.187
5. Shaoyang 2.969	45. Qinhuangdao 1.729	85. Kunming 1.167
6. Xiangyang 2.912	46. Xianyang 1.724	86. Anshan 1.137
7. Xinxiang 2.91	47. Urumqi 1.661	87. Changsha 1.131
8. Nanyang 2.755	48. Tai'an 1.656	88. Hefei 1.128
9. Xiamen 2.732	49. Rizhao 1.648	89. Heze 1.124
10. Baoji 2.436	50. Hengshui 1.644	90. Quzhou 1.123
11. Anyang 2.406	51. Puyang 1.624	91. Zhanjiang 1.117
12. Zhengzhou 2.379	52. Cangzhou 1.597	92. Sanmenxia 1.09
13. Zhumadian 2.363	53. Lanzhou 1.568	93. Zibo 1.085
14. Qingdao 2.358	54. Jining 1.553	94. Jiangmen 1.075
15. Putian 2.342	55. Hanzhong 1.537	95. Weifang 1.071
16. Wenzhou 2.247	56. Linyi 1.517	96. Shenyang 1.066
17. Lianyungang 2.197	57. Changchun 1.501	97. Shantou 1.054
18. Kaifeng 2.195	58. Nantong 1.464	98. Baotou 1.042
19. Fushun 2.118	59. Hohhot 1.461	99. Chengde 1.02
20. Shijiazhuang 2.083	60. Tangshan 1.432	100. Foshan 0.98
21. Handan 2.039	61. Quanzhou 1.425	101. Jiujiang 0.949
22. Luoyang 2.011	62. Chengdu 1.425	102. Yinchuan 0.936
23. Nanjing 2.001	63. Heyuan 1.422	103. Yichang 0.886
24. Liuzhou 2	64. Qingyuan 1.398	104. Daqing 0.879
25. Ganzhou 1.994	65. Fangchenggang 1.367	105. Fuxin 0.87
26. Shangqiu 1.991	66. Harbin 1.352	106. Zhaoqing 0.863
27. Zhoukou 1.98	67. Dongying 1.339	107. Zhoushan 0.855
28. Nanning 1.959	68. Wuhan 1.326	108. Xining 0.833
29. Weihai 1.908	69. Wuxi 1.303	109. Bengbu 0.795
30. Langfang 1.896	70. Huaian 1.291	110. Huizhou 0.789
31. Baoding 1.896	71. Yantai 1.273	111. Dezhou 0.77
32. Zhangzhou 1.872	72. Taiyuan 1.269	112. Yangjiang 0.764
33. Guilin 1.859	73. Taizhou 1.244	113. Dongguan 0.709
34. Hangzhou 1.845	74. Yingkou 1.239	114. Xuchang 0.663
35. Nanchang 1.82	75. Shaoxing 1.234	115. Maoming 0.662
36. Xingtai 1.815	76. Chongqing 1.234	116. Huzhou 0.527
37. Liaocheng 1.812	77. Ningde 1.225	117. Jiaxing 0.523
38. Shenzhen 1.809	78. Beihai 1.222	
39. Zhangjiakou 1.809	79. Jingmen 1.221	
40. Tianjin 1.801	80. Zhongshan 1.207	

Table 3: Supply Constraints and House Price

Panel	Panel A: baseline estimates		Panel B: first-stage results	
Dependent variable	Log (real house price)		Regulatory tax × Salary	Geographical constraint × Salary
Specification	(1)	(2)	(3)	(4)
	OLS	IV		
Log (real salary)	0.1840***	0.1714***	0.1623	-1.0402***
	(0.0398)	(0.0570)	(0.1360)	(0.1352)
Regulatory tax × log (real salary)	0.0247***	0.1097***		
	(0.0079)	(0.0353)		
Geographical constraint × log (real salary)	0.0749***	0.0723**		
	(0.0128)	(0.0324)		
Local tax enforcement × log (real salary)			0.3390***	0.3251***
			(0.0541)	(0.0438)
Number of dialects × log (real salary)			0.1059***	-0.0808***
			(0.0317)	(0.0179)
Distance to major lakes × log (real salary)			-0.0214	-0.1909***
			(0.0319)	(0.0237)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
Time period	2005-2018		2005-2018	
Number of cities	105		105	
N	1444	1444	1444	1444
R^2	0.9707		0.9989	0.9994
Kleibergen-Paap rk Wald F-statistic		12.52		

Notes: Both supply constraints measures are standardised. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively. Robust standard errors in parentheses.

Table 4: Effect of Shifts in Real Salary on House Price in Average Chinese City (Counterfactual Outcomes)

Variable	Value in 2005	Value in 2018	SD	Max	Min
Predicted house price (RMB/m^2)	2381	5260	2735	15454	1921
Panel A: supply constraints and demand shock set to zero sequentially					
Predicted without any regulatory tax	2381	4174	2170	12262	1524
- and share developed set to zero	2381	4001	2081	11756	1461
- and independent effect of salary removed	2381	3378	1756	9923	1233
Panel B: supply constraints lowered by 1 std. dev. and demand shock set to zero sequentially					
Predicted with regulatory tax lowered by 1 std. dev.	2381	4720	2454	13866	1723
- and share developed lowered by 1 std. dev.	2381	4394	2285	12908	1604
- and independent effect of salary removed	2381	3709	1928	10895	1354
Panel C: supply constraints set to zero separately					
Predicted without any regulatory tax	2381	4174	2170	12262	1524
Predicted with % developed set to zero	2381	5043	2622	14816	1841
Panel D: supply constraints lowered by 1 std. dev. separately					
Predicted with regulatory tax lowered by 1 std. dev.	2381	4720	2454	13866	1723
Predicted with % developed lowered by 1 std. dev.	2381	4897	2546	14387	1788

Figures

Fig. 1: Housing Price in Major Chinese Cities

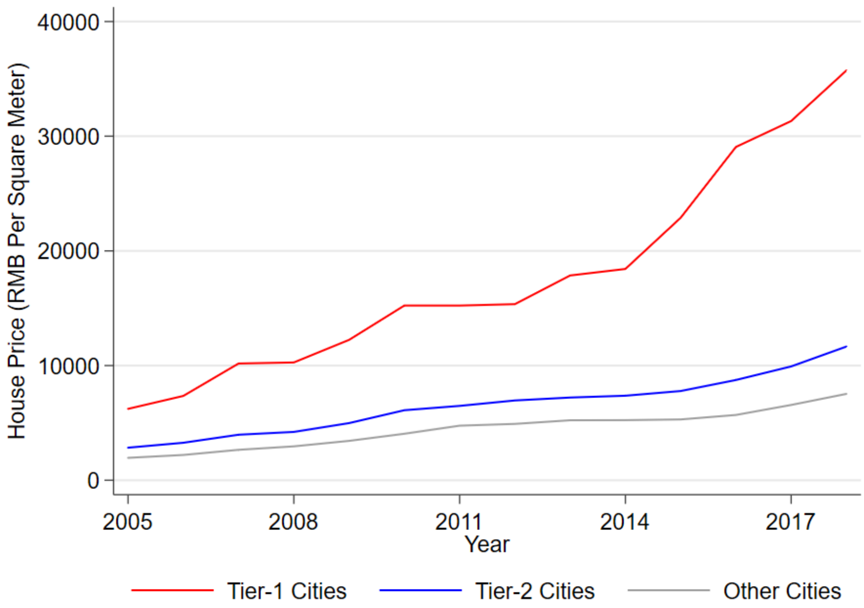


Fig. 2: House Price, Average Salary, and Construction Cost

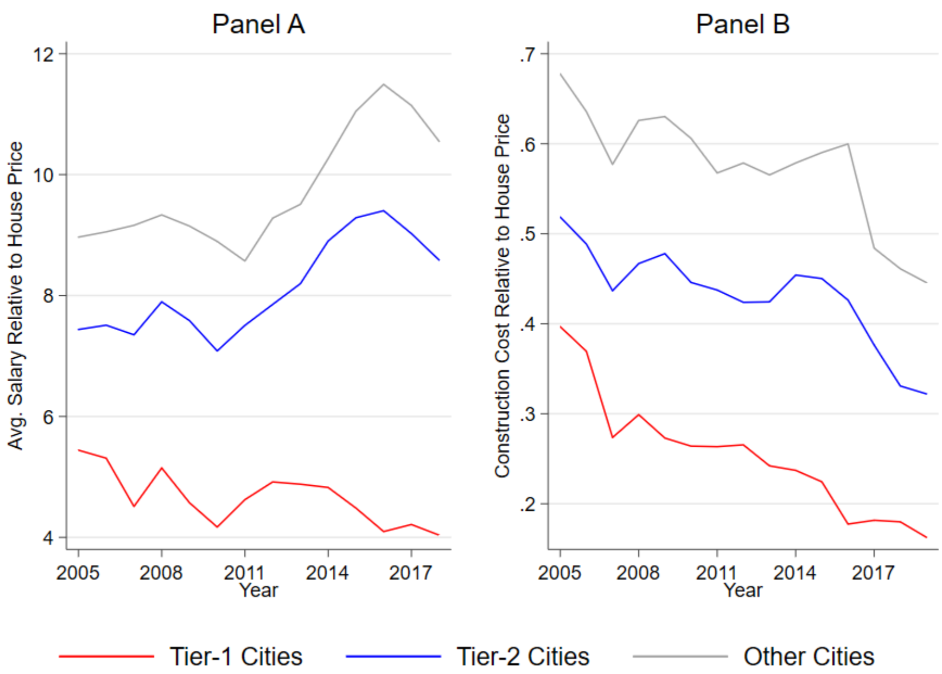


Fig. 3: Land Supply in Major Chinese Cities

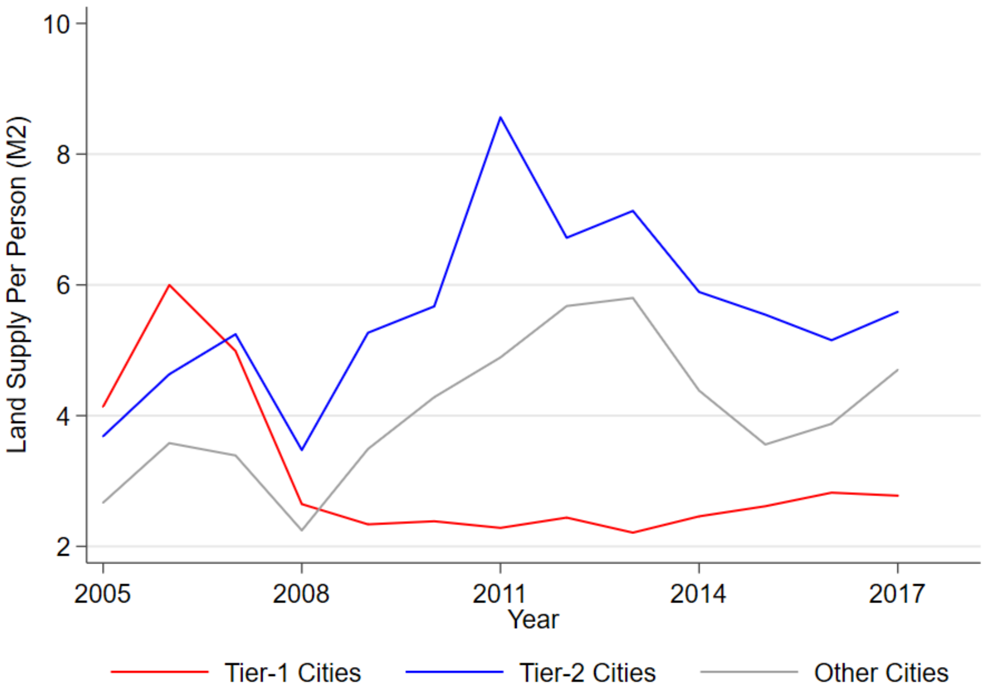


Fig. 4: Regulatory Tax Estimate at the Project Level

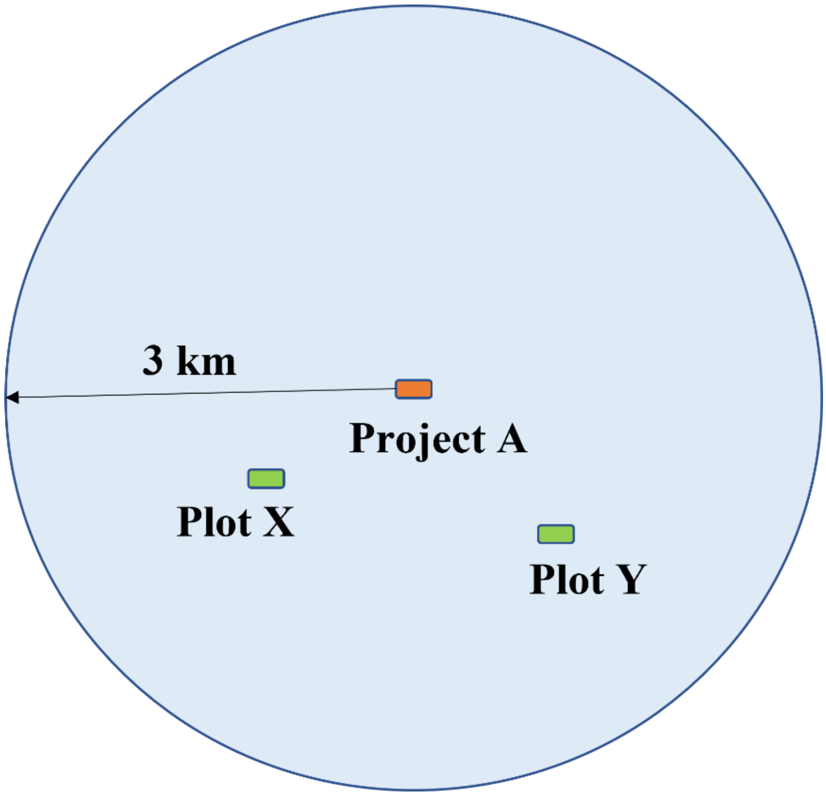


Fig. 5: Histograms

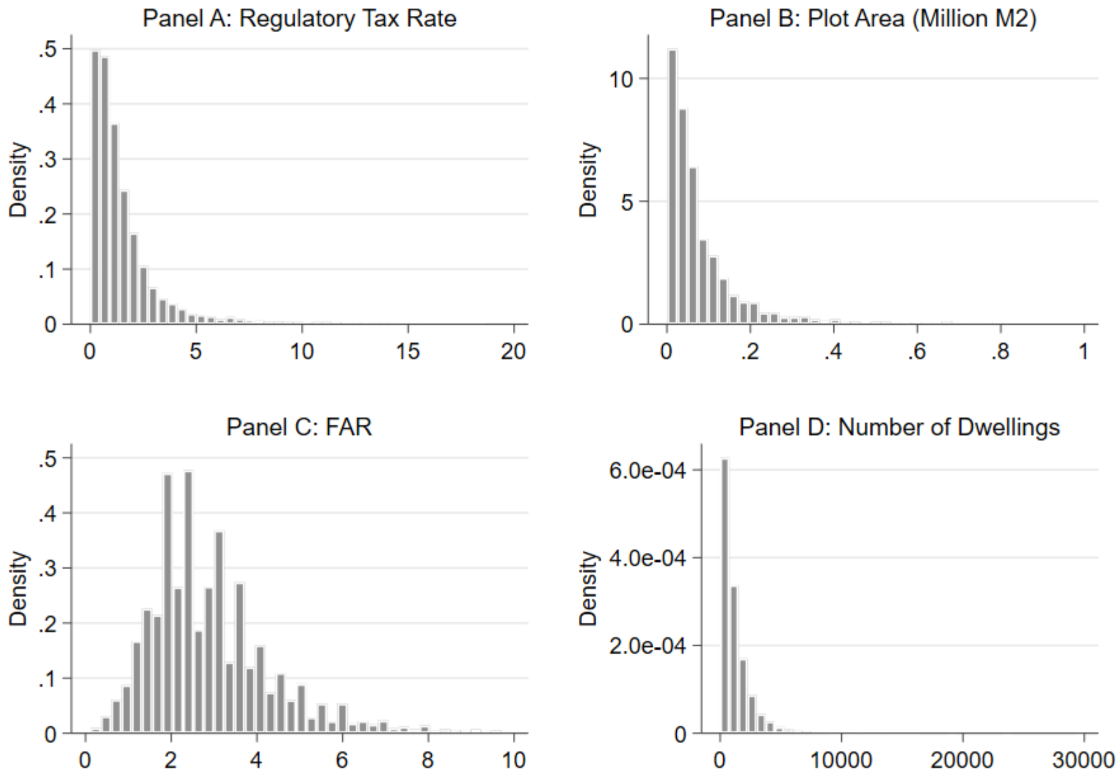


Fig. 6: House Price Decomposition in Tier-1 Cities and Jiaxing

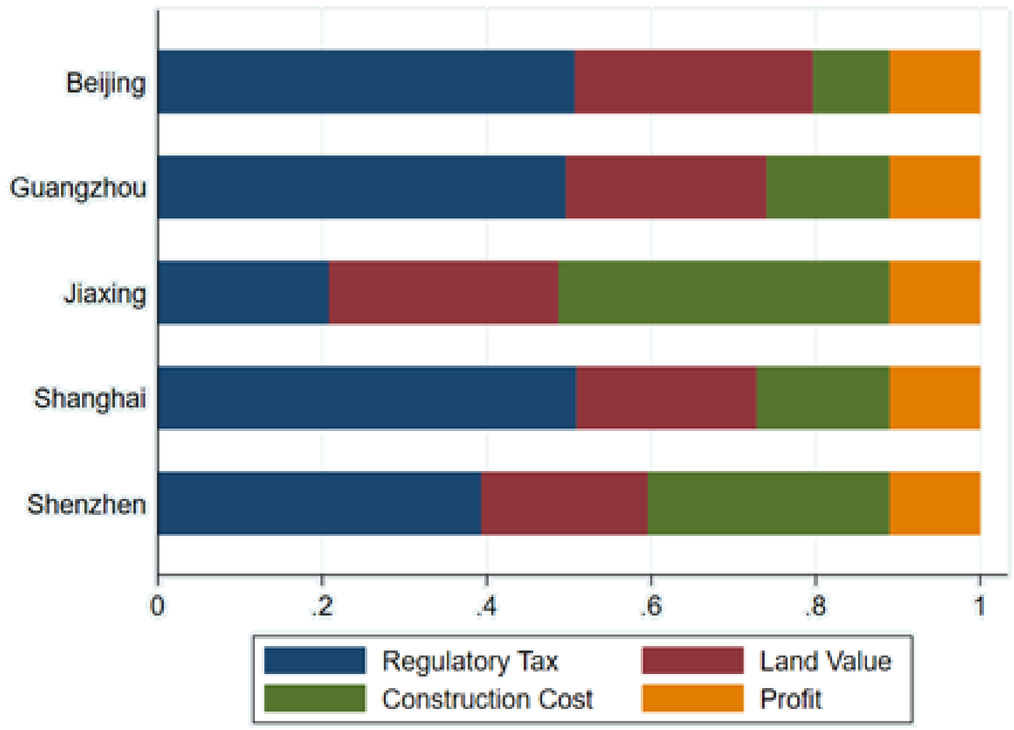


Fig. 7: House Price Decomposition for Different Tiers of Cities

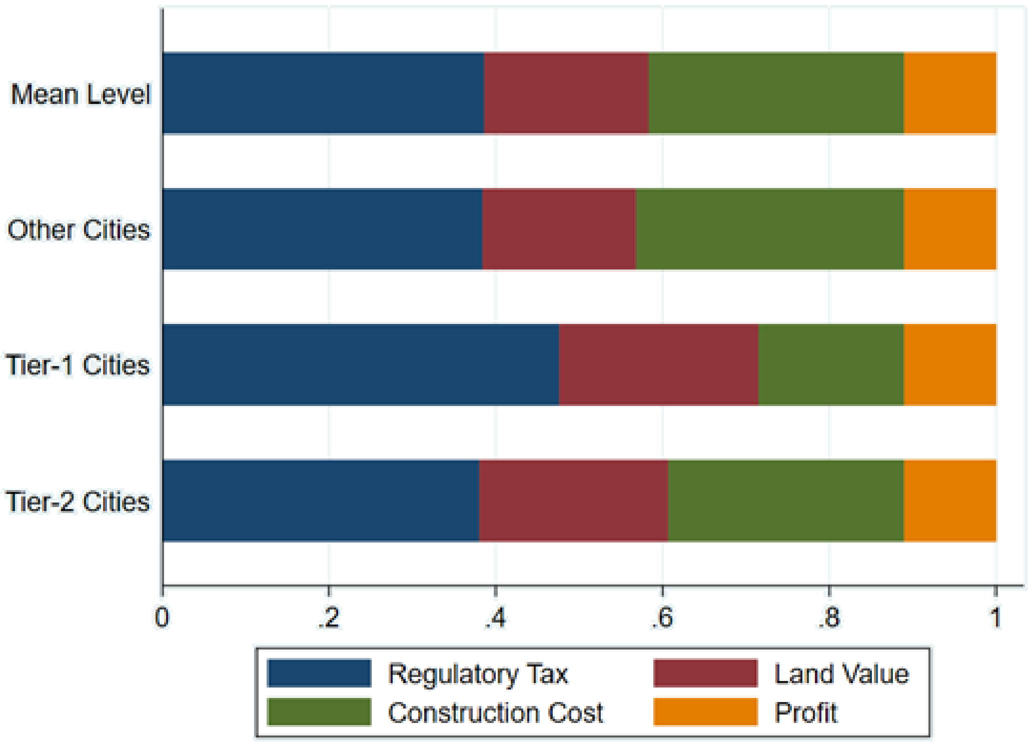


Fig. 8: Regulatory Tax Rate

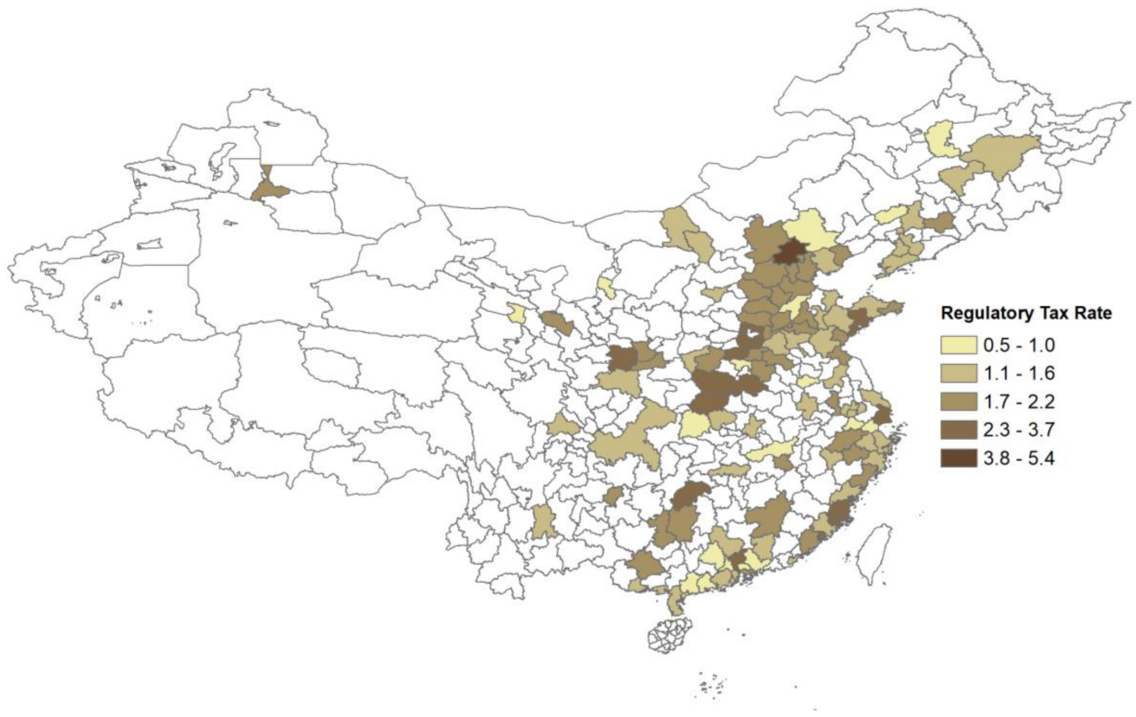


Fig. 9: Scatter Plots

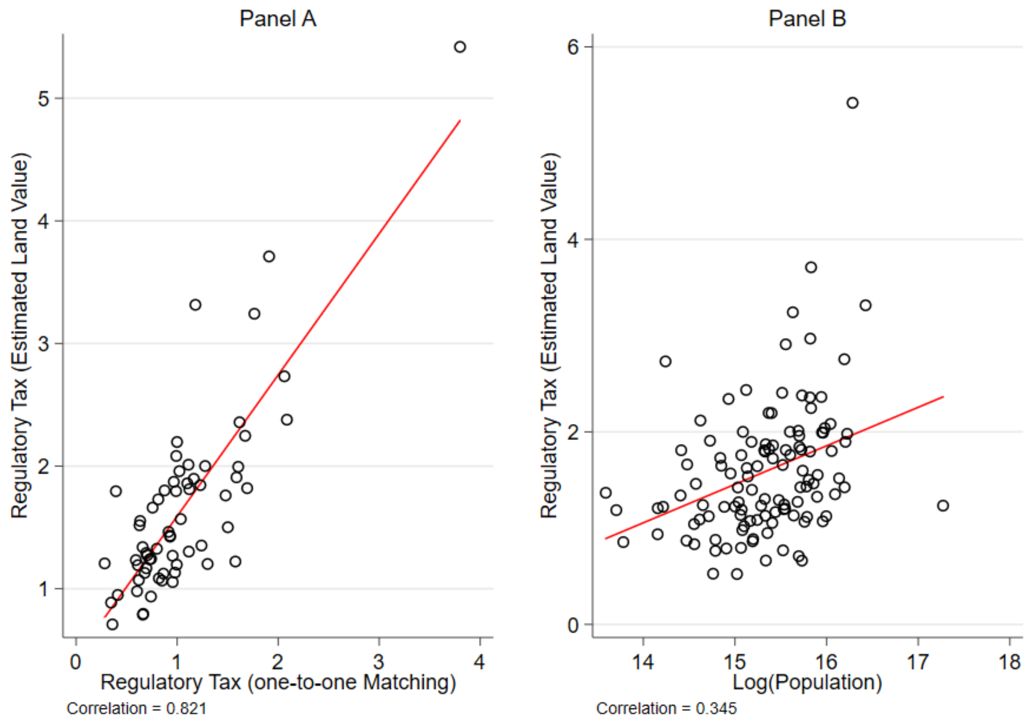


Fig. 10: Demand Shock and Housing Supply Constraints

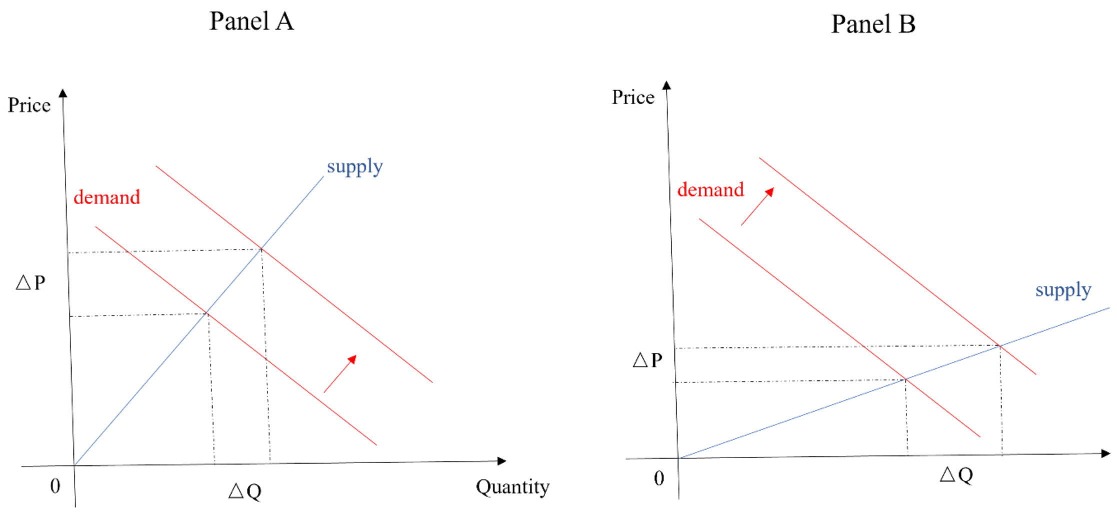


Fig. 11: Impact of Removing Supply Constraints on House Prices in Average Chinese City

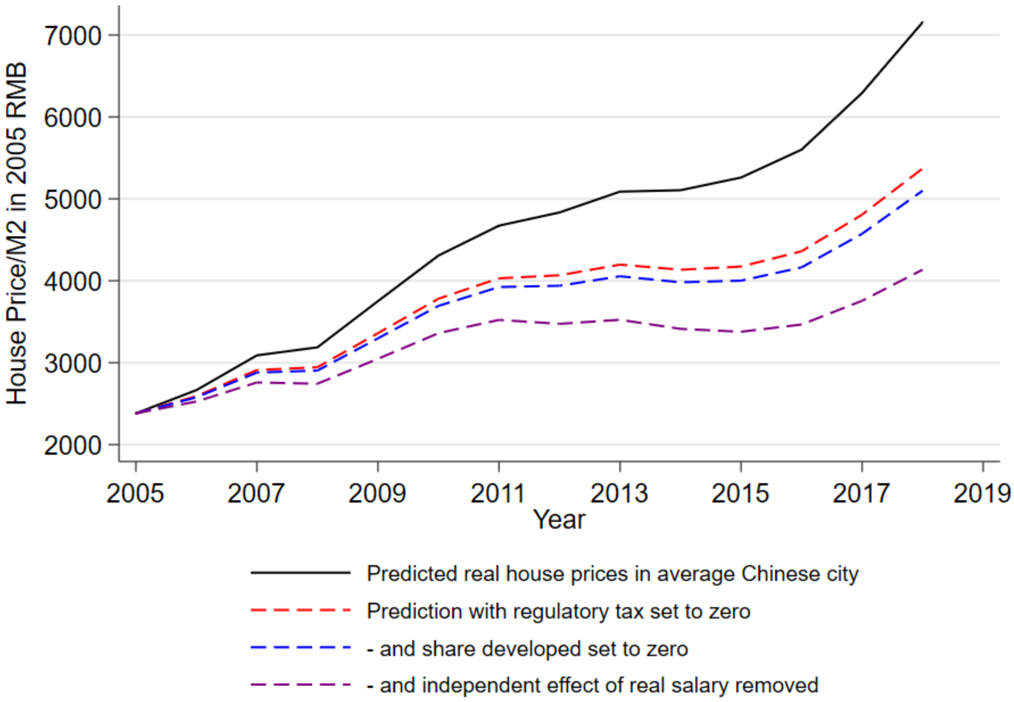


Fig. 12: Impact of Reducing Supply Constraints on House Prices in Average Chinese City

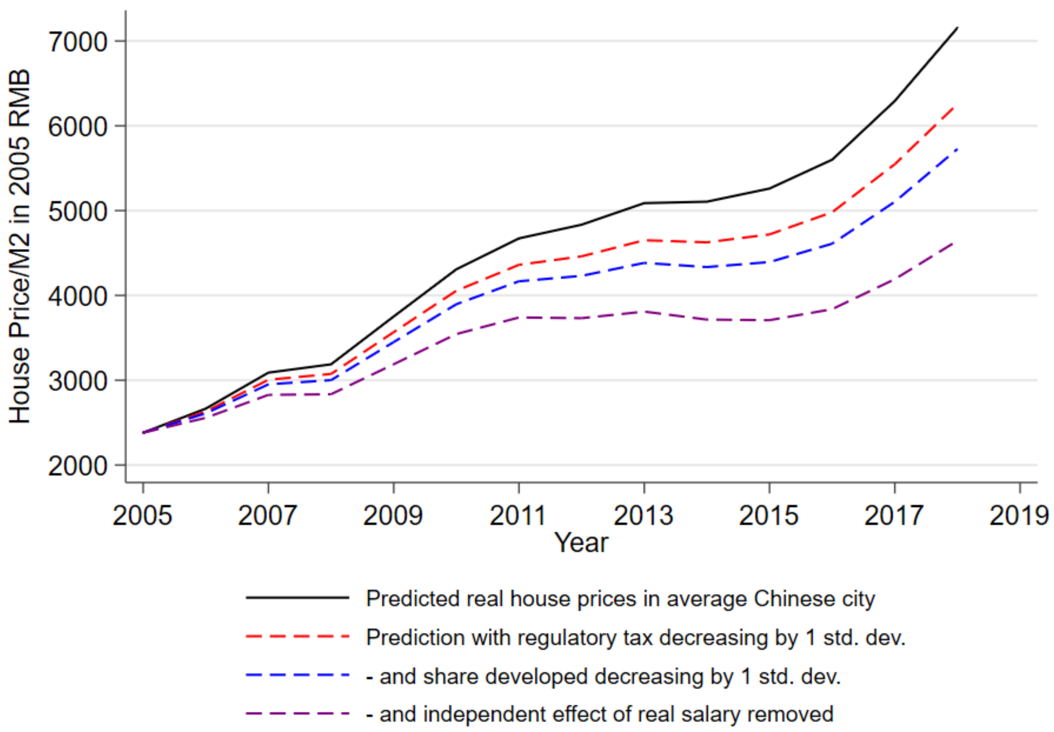
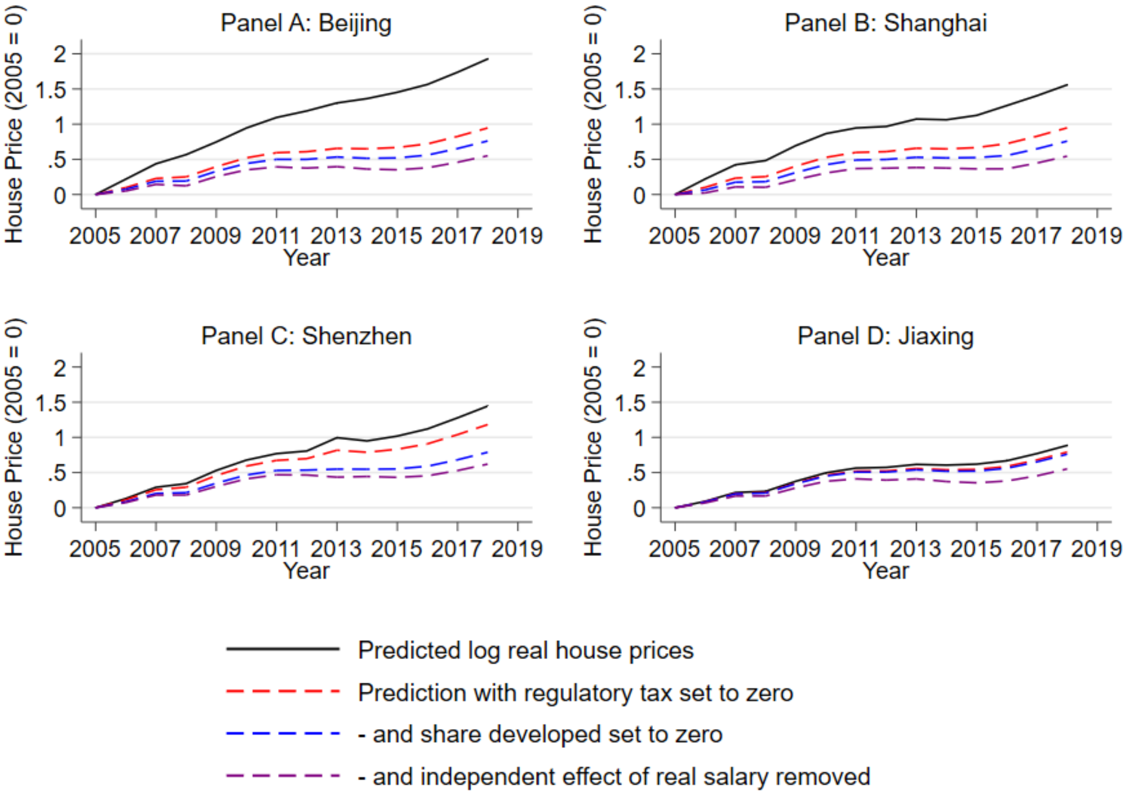


Fig. 13: Predicted Log of Real House Prices in Selected Cities



Appendices

Appendix A: Appendix Tables

Table A1: Development Cost Breakdown

Code	Cost item
1	Land acquisition fees
1.1	Land acquisition fees
1.2	Land premium
1.3	Land taxes and fees
1.4	Other expenses
2	Early-stage engineering costs
2.1	Feasibility study
2.2	Survey and measurement
2.3	Planning and design
2.4	Land clearing and utility connection
2.5	Temporary facility
2.6	Administrative fees
2.7	Other expenses
3	Construction fees
3.1	Foundation
3.2	Main project
3.3	Decoration
3.4	Mechanical and electrical equipment installation
4	Infrastructure fees
4.1	Municipal projects
4.2	Landscape
5	Public facilities fees
5.1	Public construction costs
5.2	Other expenses

Table A2: Construction, Land Supply, and Supply Constraints

Dependent variable	Log (land supply)		Log (construction)	
Specification	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
Log (real salary)	1.2839***	1.0263***	0.8253***	0.8130***
	(0.1976)	(0.2663)	(0.1565)	(0.2061)
Regulatory tax × log (real salary)	0.0046	0.2190	-0.0612*	-0.4804***
	(0.0444)	(0.1667)	(0.0318)	(0.1178)
Geographical constraint × log (real salary)	-0.2073***	-0.4153***	-0.1656***	-0.0277
	(0.0716)	(0.1547)	(0.0445)	(0.0953)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
Time period	2005-2017		2005-2018	
Number of cities	105		103	
N	1332	1332	916	914
R^2	0.7453		0.9489	
Kleibergen-Paap rk Wald F-statistic		12.24		8

Notes: Both supply constraints measures are standardised. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively. Robust standard errors in parentheses.

Table A3: House Price and Supply Constraint (Local Labour Demand Shock)

Panel	Panel A: baseline estimates		Panel B: first-stage results	
Dependent variable	Log (real house price)		Regulatory tax × LLDS	Geographical constraint × LLDS
Specification	(1)	(2)	(3)	(4)
	OLS	IV		
LLDS	-0.2575***	-0.2357**	-0.3150*	-0.1231
	(0.0979)	(0.1051)	(0.1881)	(0.1318)
Regulatory tax × LLDS	0.0638***	0.1907***		
	(0.0168)	(0.0627)		
Geographical constraint × LLDS	0.0829***	0.0968**		
	(0.0200)	(0.0409)		
Local tax enforcement × LLDS			0.3197***	0.4245***
			(0.0327)	(0.0509)
Number of dialects × LLDS			0.1440***	-0.1283***
			(0.0242)	(0.0203)
Distance to major lakes × LLDS			-0.0401*	-0.2878***
			(0.0240)	(0.0288)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
Time period	2005-2018		2005-2018	
Number of cities	104		104	
N	1444	1444	1444	1444
R^2	0.9701		0.9990	0.9984
Kleibergen-Paap rk Wald F-statistic		33.87		

Notes: Both supply constraints measures are standardised. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively. Robust standard errors in parentheses.

Table A4: Alternative Estimates of Regulatory Tax Rate

Dependent variable	Log (house price)		Log (house price)	
Specification	(1)	(2)	(1)	(2)
	OLS	IV	OLS	IV
	One-to-one matching		Perfect competition	
Log (real salary)	0.0970*	0.2181***	0.1890***	0.1823***
	(0.0535)	(0.0841)	(0.0396)	(0.0569)
Regulatory tax × log (real salary)	0.0316***	0.0664**	0.0240***	0.1050***
	(0.0119)	(0.0271)	(0.0078)	(0.0355)
Geographical constraint × log (real salary)	0.0531***	0.1340***	0.0744***	0.0773**
	(0.0131)	(0.0430)	(0.0128)	(0.0317)
Year FEs	Yes	Yes	Yes	Yes
City FEs	Yes	Yes	Yes	Yes
Time period	2005-2018		2005-2018	
Number of cities	60		107	
N	831	831	1472	1472
R^2	0.9652		0.9708	
Kleibergen-Paap rk Wald F-statistic		17.3		12

Notes: Both supply constraints measures are standardised. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively. Robust standard errors in parentheses.

Table A5: Drop Tier-1 Cities and Municipalities

Dependent variable	Log (house price)	
Specification	(1)	(2)
	OLS	IV
Log (real salary)	0.1700***	0.1434**
	(0.0401)	(0.0685)
Regulatory tax × log (real salary)	0.0232**	0.1864**
	(0.0099)	(0.0753)
Geographical constraint × log (real salary)	0.0499***	0.0820*
	(0.0124)	(0.0472)
Year FEs	Yes	Yes
City FEs	Yes	Yes
Time period	2005-2018	
Number of cities	99	
N	1362	1362
R^2	0.9653	
Kleibergen-Paap rk Wald F-statistic		7.2

Notes: Both supply constraints measures are standardised. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively. Robust standard errors in parentheses.

Appendix B: Appendix Figures

Fig. B1: Share of Developed Land in 2005

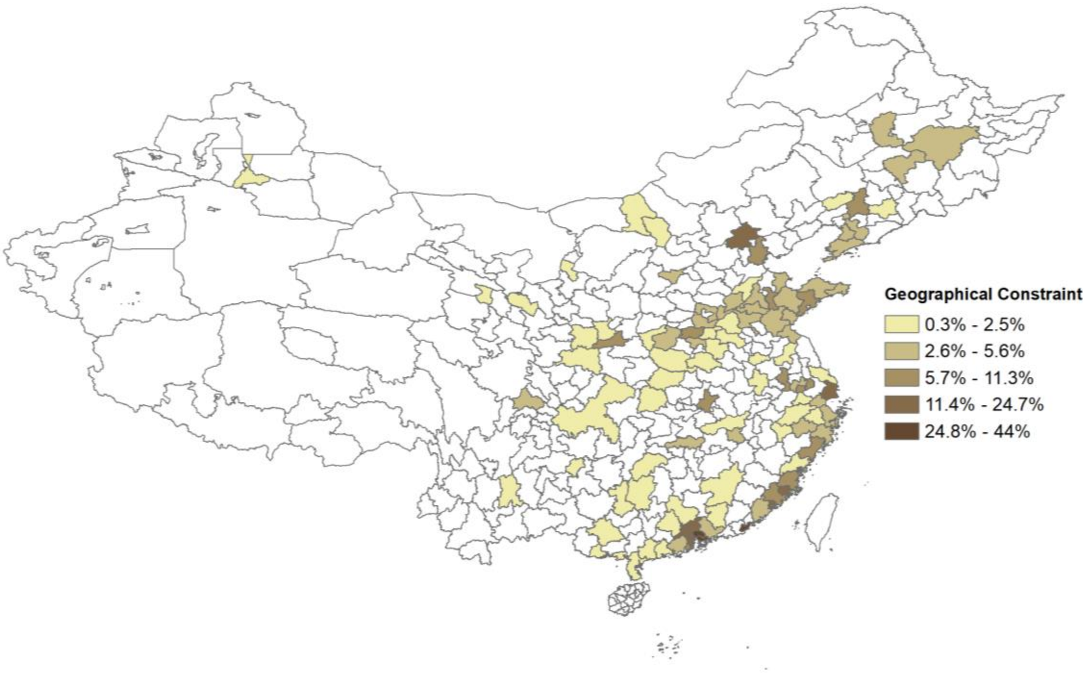


Fig. B2: Number of Dialects

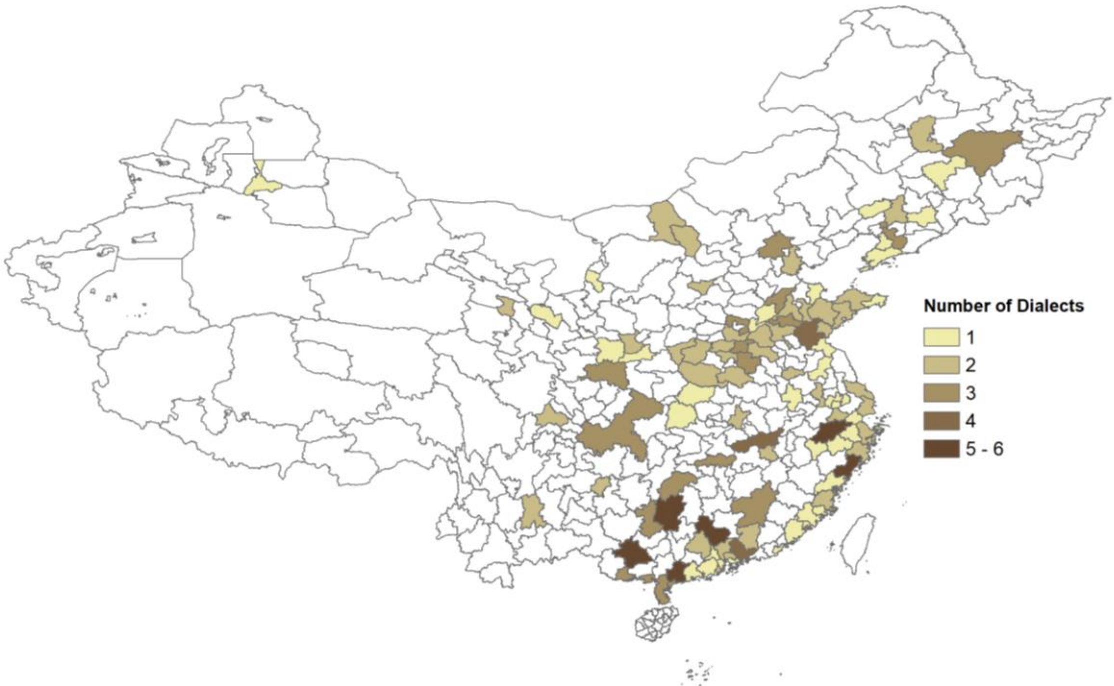


Fig. B3: Major Lakes and Rivers in 1820

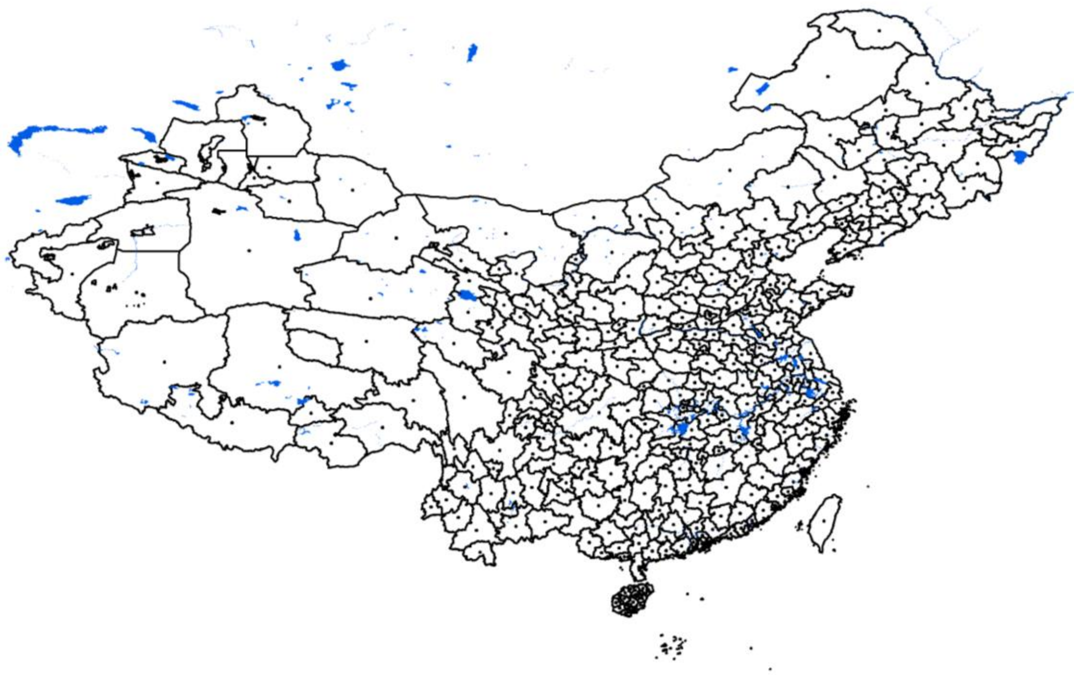


Fig. B4: Industry Index in China

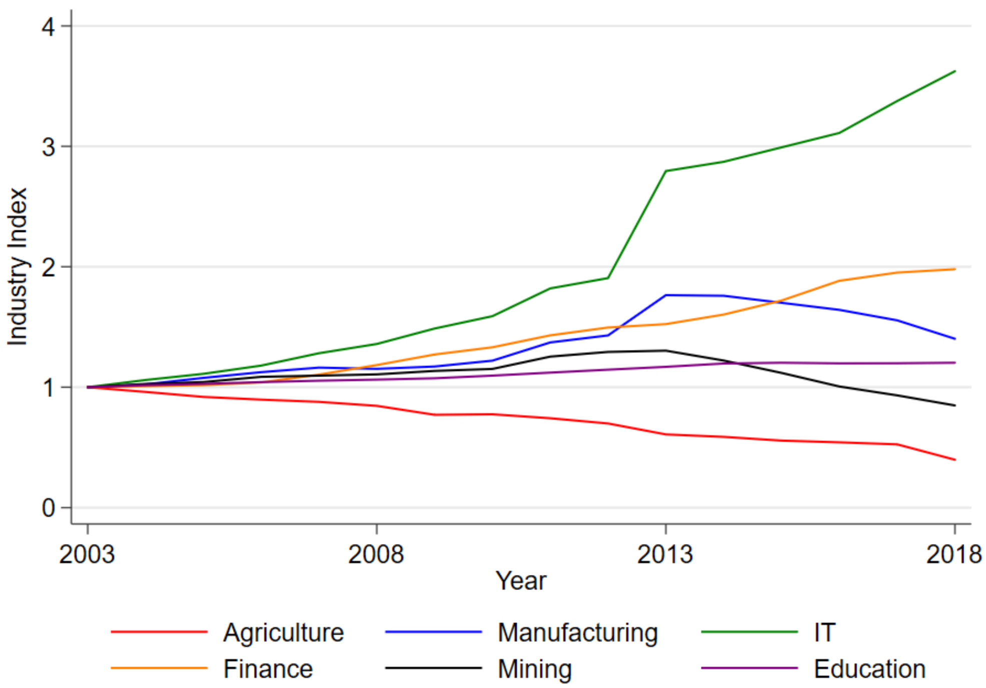
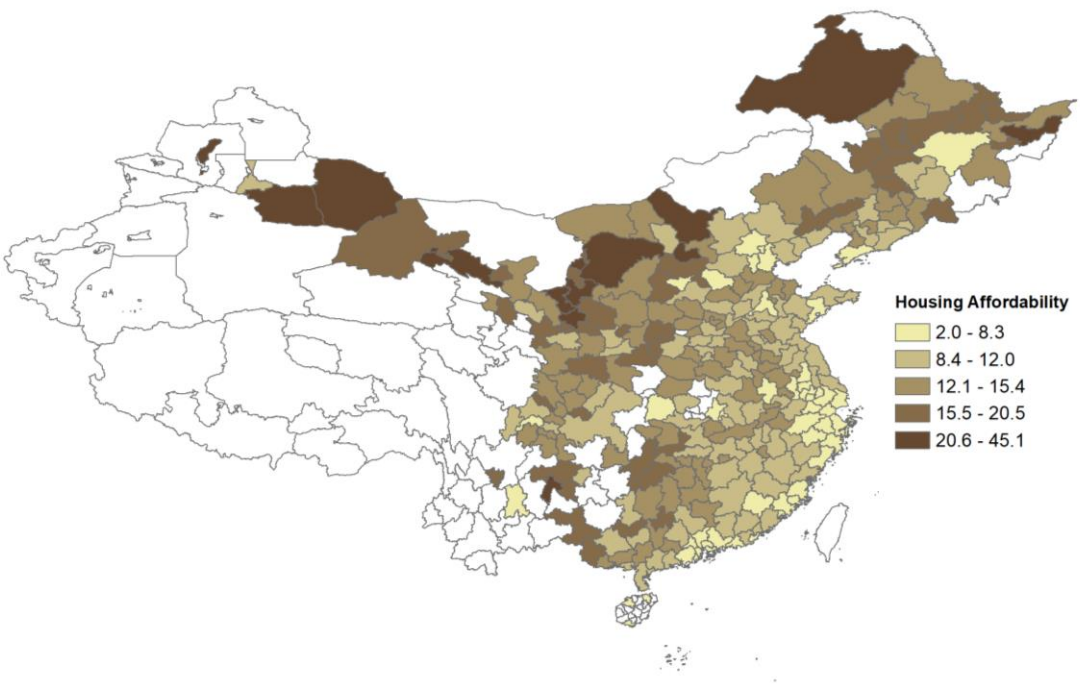


Fig. B5: Housing Affordability in China in 2018



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