

# Housing Policy with Home-Biased Landlords: Evidence from French Rental Markets

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## Abstract

Tax incentives are widely used to encourage the construction of rental dwellings, and reduce housing costs for low-income households. This paper examines an important source of local variation in the impact of these policies: the home-biased supply of capital from potential landlords. Building upon administrative data covering the universe of dwellings in France, I document that private investors own undiversified and lumpy amounts of rental housing, face substantial return frictions across projects and locations, and strongly prefer properties close to their own home. A spatial equilibrium framework combined with a frictional portfolio choice rationalizes these features. Because spatial frictions reduce the private return to more remote rental investments, mobile agents, who tend to be renters, agglomerate near the residence of affluent and immobile owners. The model also predicts that home bias in the supply of capital regulates the heterogeneity in housing market responses to place-based subsidies. Exploiting quasi-experimental evidence from a location-specific investment tax credit targeted at individual landlords in France, the *Pinel* law, I evidence its substantial causal impact on dwelling sales and new construction. In the medium-run, subsidies expand the local housing stock with limited crowding-out, inducing substantial in-migration of renters and population growth. Out-of-town individual investor involvement rises in treated cities, and the policy has stronger effects in locations more open to outside capital. The incomplete capitalization of the incentive in new unit prices confirms that under imperfect capital mobility, landlords bear part of the incidence of local subsidies in the form of higher net returns. Therefore, landlord home bias is a key factor in the allocation of capital and people across space, and determines the efficiency and distributional effects of place-based policies.

**JEL codes:** E22, G11, H31, R31, R38.

**Keywords:** housing supply, investment, landlords, spatial equilibrium, tax subsidies.

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*The landlords, like all other men, love to reap where they never sowed.*

Adam Smith, *An Inquiry into the Nature and Causes of the Wealth of Nations*, I, 6.

## 1 Introduction

A limited supply of housing constrains the growth of cities (Glaeser and Gyourko, 2018). In particular, the scarcity of rental accommodation – a key driver of spatial mobility – severely limits migration towards desirable or productive areas. In response, governments in advanced economies have devised a wide array of place-based interventions to alleviate housing burdens for renters in expensive real estate markets. These regulatory and fiscal tools must balance two conflicting goals: making rental housing affordable for tenants, while letting its provision remain profitable to landlords. Navigating this trade-off, in turn, requires understanding the constraints that govern investment behavior across local housing markets. Do financial and distance barriers affect whether individual savers decide to become landlords, and where? To what extent do tax incentives shape the geography of new housing investment? And how do spatial frictions determine the way these policies impact the portfolio allocation of investors, the supply of housing, and the mobility of renters?

Most studies of the forces that lead economic activity to cluster in space document channels operating in the markets for labor (Marshall, 1890), knowledge (Jacobs, 1969), or final goods (Krugman, 1991). Paradoxically, on the other hand, public policies shaping the economics of cities concentrate on housing investment, where academic research generally abstracts from space and assumes a perfectly elastic supply of “absentee” capital. This paper provides a possible reconciliation between the policy focus on housing production, and the theoretical argument for endogenous agglomeration. It illustrates empirically and conceptually that landlord ownership decisions are biased towards rental properties close to their own residence. Notwithstanding Adam Smith’s view, landlords love to reap precisely *where* they sowed. The home bias of these immobile individual savers, who supply imperfectly elastic capital across space, affects the provision of rental housing. It makes it relatively cheaper for tenants to agglomerate in cities with high proximity to affluent savers – a “market access” term, in the terminology of Redding and Rossi-Hansberg (2017), that operates through frictional capital markets. Critically, spatial frictions also shape the response of housing markets to tax incentives. While individual residential investment, like business capital formation (Zwick and Mahon, 2017), responds to tax policies, its specific spatial pattern has implications for the mechanisms and heterogeneity at play. Using quasi-experimental evidence from a French investment tax credit, I demonstrate how, consistent with a stylized model, the geography of home-biased landlords influences the heterogeneous effects of place-based rental subsidies across eligible locations, and regulates their impact on the mobility of capital and people in the economy.

To investigate constraints to the supply of rental services by individual savers, and analyze the dynamics of production, sales, and occupancy for the rental housing stock, I assemble register data on the universe of close to thirty million primary residences in France. I also avail myself of exhaustive administrative records for

all real estate transactions, building permits requests, and individual landlord tax returns, along with detailed data on land use, social housing, mobility, and employment, covering France's 35,000 municipalities over the 2010-2020 decade. The paper then proceeds in three main steps.

I begin by documenting three facts illustrating that rental housing investment is a highly frictional venture. Reminiscent of common paradoxes in international finance (Feldstein and Horioka, 1980; Lucas, 1990; French and Poterba, 1991), rental capital flows within a country exhibit significant spatial segmentation, evidenced by three regularities: portfolio lumpiness, return dispersion, and local bias in ownership.

First, wealth survey microdata illustrate that rental property is mostly held in lumpy quantities by undiversified and mostly immobile landlords. Far from the ideal view of a common diversified housing market portfolio, each of them tends to operate a small number of units, a limited scale suggestive of physical indivisibility and financial constraints. Next, I show there exists substantial dispersion in rent-to-price ratios, both across cities and among individual units within narrow locations. That variation in capitalization rates differs from their hypothetical equalization to a unique user cost, as expected in integrated capital markets. While it could stem from differences in risk or expected capital gains, the additional fact that rent-to-price ratios decline with the distance of a location from potential investors suggests that heterogeneous access to investors matters in the market for residential capital. Lastly, an analysis of more than seven million rental properties in France demonstrates that investment linkages between cities (the probability of ownership of rental assets in one of them by investors from the other) decay rapidly with their physical distance and estimated social connections. The geography of rental asset ownership displays a substantial "housing home bias" of landlords for dwellings near their own residence location.

Together, these regularities make investment frictions essential in shaping the allocation of capital across markets, and, in turn, of mobile workers across space. Building upon and rationalizing these facts, my second contribution is a stylized model where barriers to cross-city capital mobility constrain new housing production, and influence its response to tax policies. While the imperfect mobility of *workers* (Kline and Moretti, 2014a) and *firms* (Fajgelbaum et al., 2019) has been the topic of a long spatial literature, distance barriers to the allocation of *housing capital* by individual investors remain unexplored.

In the model, renters, as in a standard Rosen (1979)-Roback (1982) framework, move across locations to arbitrage real wage differentials. Housing producers build new units on local land, but do not lease them directly; rather, they sell them outright to investors, who are the sole providers of rental services. I depart, however, from the standard assumption that deep-pocketed, well-diversified, and "absentee" landlords require a uniform rate of return on their rental properties, a perspective under which the free flow of capital would equalize rent-price ratios across locations. Instead, these dwellings are owned by overlapping generations of spatially static individual savers, whose location-specific rental asset demand reflects relative returns, operating costs, and idiosyncratic frictions or preferences for operating specific residential projects.

Because these subjective investment frictions increase with distance, landlords prefer to operate rental

dwelling close to their own residence, willingly accepting lower gross returns on nearby properties. Their lumpy investment decisions and responsiveness to return differentials shape housing affordability in destination markets. As in the data, the ownership of the rental stock exhibits a model-implied gravity structure, and areas located close to an ample supply of investible capital have lower rent-to-price ratios, due to their higher “investor portfolio access”. Alongside long-studied local externalities (Duranton and Puga, 2004), the capital-intensive but frictional production of rental services thus creates a novel agglomeration force for the location choice of mobile workers – who are overwhelmingly tenants. This frictional investment behavior also pins down spatial heterogeneity in the tax responsiveness of rental housing supply: locations more open to outside capital draw from a broader and more distant pool of individual investors when landlord subsidies are implemented. With potential landlords reluctant to invest far from home, the market for rental assets is segmented, and the upwards-sloping supply of capital less than perfectly elastic across space: the incidence of place-based housing subsidies partly benefits infra-marginal savers in the form of higher returns.

In the third part of the paper, I present direct quasi-experimental evidence that, consistent with the model, individual landlords not only respond sharply to housing tax policies, but also provide an imperfectly elastic capital supply across space. To quantify the causal response of landlord behavior to place-based incentives, I study a French tax credit targeted at the provision of new rental housing by individual investors. The scheme, known as the *Pinel* law, offers taxpayers a substantial personal income tax reduction (of up to 21 percent of the purchase price) if they buy and lease newly built properties in targeted high-cost areas, conditional on respecting affordable rent levels and tenant income ceilings. I take advantage of variation across comparable municipalities in the place-based eligibility of buy-to-let subsidies after their introduction in 2014, and of their partial removal in 2018, when the *Pinel* incentive was discontinued in some cities as a byproduct of nationwide budget cuts. Leveraging these reforms, I combine difference-in-differences around entry and exit for similar treated and control locations, to test model predictions on the consequences of place-based subsidies.

First, tax policy causally affects the allocation of individual *investment* across space. Developer sales of eligible newly built dwellings to individual investors increase by 18-20 percent each quarter in targeted locations for the full duration of the subsidy (4 years), with the effect operating at both the intensive and extensive margins. The implied user cost elasticity of investment ranges from 1.3 to 2, depending on an estimated net-of-rent-discounts present value of the incentive between 9 and 15 percent. Far from merely accelerating sales of existing developer inventory, the policy triggers entirely new construction projects, increasing local housing starts and land demand. The composition of dwellings also shifts, as the share of multi-family buildings and social housing increases, and urban sprawl rises. Going beyond the difference-in-differences design, in the longer-run, nationwide rental property purchases closely co-move with the time-varying generosity of government subsidies, as predicted by the model. Total new home sales also display stark time-bunching in anticipation of eligibility deadlines for buy-to-let tax credits, and abnormal end-of-fiscal-year spikes, hinting at the salience of taxation for aggregate housing investment.

Next, I provide direct evidence that the effects are consistent with the spatial bias mechanism of the concep-

tual framework. First, in contrast to common findings for real estate subsidies, but in line with an imperfectly elastic capital supply across space, there is only incomplete capitalization of the incentive in new unit prices. The estimated price effects only amount to between one third and one half of the net-of-rent-discount present value of the landlord subsidy. The policy thus raises post-tax returns to rental property investment in treated locations to attract additional capital. Next, as predicted by the model's extensive margin of portfolio reallocation, the scheme operates by drawing in more distant marginal investors to treated locations, somewhat offsetting home bias in rental investment. Finally, I show evidence that the scheme displays stronger effects in locations more open to outside ownership, confirming that the degree of *capital* mobility has far-reaching consequences for the provision of housing across places. The policy results in a 2 percent long-lived expansion in the local housing stock, with a larger effect in locations characterized by a more spatially dispersed ownership.

Last, the portfolio choice of investors entails a downstream reshuffling of *people*: spatial capital reallocation induces labor mobility. Contrary to the existing literature on housing policy, I find no evidence of crowding-out of non-subsidized housing. As a consequence of this limited crowding-out, in the medium-run, the scheme triggers a positive response of geographic mobility towards towns eligible to the individual landlord subsidy. It leads to increases in local population and employment, mediated by a sharp upwards jump in inwards migration (notably of middle- and lower-income tenants) towards treated cities.

**Related literature** This paper evidences the relationship between the spatial allocation of economic activity, the provision of housing, and the geography of wealth. It builds on three main strands of investigation: on the implications of housing supply for the spatial economy; on the heterogeneous consequences of tax incentives for investment behavior; and on fiscal policies in housing markets.

First, I present a new channel for the role of housing in the spatial allocation of economic activity, operating through capital market frictions. Ganong and Shoag (2017) and Glaeser and Gyourko (2018) show that zoning regulations limiting local housing supply elasticities constrain mobility and growth, while Monte, Redding, and Rossi-Hansberg (2018) demonstrate that openness to commuting can circumvent a limited local land supply. In this literature, however, the decisions of "absentee" landlords generally take a back seat under the assumption of perfect capital mobility. I provide evidence on the specific role of private rental provision for triggering labor mobility, and explore how capital market fluidity across space matters for rental affordability. Recent work on rental housing implies that landlord heterogeneity (Greenwald and Guren, 2019) and access to finance (Gete and Reher, 2018; Reher, 2021) may generate aggregate supply curves that are upwards-sloping in returns. In my paper, systematic geographic variation in landlord user costs makes the supply of rental capital depend on physical proximity to investors, potentially misallocating the provision of dwellings.<sup>1</sup> This local bias also has distributional consequences. Due to the spatial frictions in residential investment, location-specific subsidies involve trade-offs between redistribution and incentives across places.<sup>2</sup> In the same way

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<sup>1</sup>The spatial segmentation of housing ownership that I document meets a sizable body of evidence on international (Portes, Rey, and Oh, 2001) and domestic (Coval and Moskowitz, 1999; Hong, Kubik, and Stein, 2008; Chen et al., 2010) home bias in asset holdings. The consequences of local bias for rental affordability also relate to the hypothesis that proximity to capital (Kuchler et al., forthcoming), or limited local competition in asset markets (Hong, Kubik, and Stein, 2008), have downstream effects on returns.

<sup>2</sup>Other studies, e.g. Favilukis, Mabilie, and Van Nieuwerburgh (2019), show the potential for redistribution through housing policies.

that limited *labor* mobility makes place-targeted measures fall upon infra-marginal *workers* (Kline and Moretti, 2014b; Gaubert, Kline, and Yagan, 2021), here, frictions to the allocation of *capital* across space entail a partial incidence of place-based policies on nearby *savers*.

Second, I study how frictions alter the sensitivity of investment to tax policy, a public and corporate finance question dating back to Hall and Jorgenson (1967). Zwick and Mahon (2017) and Ohrn (2018) exploit industry variation in depreciation allowances to identify the tax responsiveness of firm capital formation. Using quasi-experimental variation in *legal eligibility* (rather than differential exposure) across *locations* (rather than sectors), this paper estimates the response of *residential* investment (rather than business equipment) by *individual savers* (rather than firms).<sup>3</sup> Criscuolo et al. (2019) and Suárez Serrato and Zidar (2016) also analyze spatially varying tax incentives, but focus on firm capital and labor demand.<sup>4</sup> In this literature, heterogeneous corporate responses to taxation provide indirect evidence of firm-level financial constraints. Here, by treating landlord investment as a geographically differentiated choice, I demonstrate that heterogeneous responses to tax credits across places inform us on spatial frictions in individual portfolio allocation decisions.

Finally, I contribute to a growing body of work exploring the causal impact of government involvement in real estate markets, heretofore mostly dedicated to housing *consumption* – rather than *investment*.<sup>5</sup> The buy-to-rent subsidies I study provide tax rebates to investors in rental dwellings, often conditional on affordability criteria. In spite of the growing use – and budget costs – of subsidized privately-owned rental property, its causal effect on investment behavior, spatial asset ownership, and tenant mobility remains largely unknown, along with its distributional consequences. In the United States, the Low-Income Housing Tax Credit (LIHTC), whereby the federal government shoulders part of the cost of affordable housing construction by corporate developers, has attracted attention as its fiscal burden rose in recent years (Baum-Snow and Marion, 2009; Diamond and McQuade, 2019). The French *Pinel* law shares similarities with the LIHTC: both provide tax benefits to newly built dwellings in exchange for rent moderation. However, relative to the LIHTC, the *Pinel*'s open-ended nature, the nationally pre-determined eligibility of locations, and its targeting of individual taxpayers, further our understanding of residential investment responses, capital market segmentation in housing production, and agglomeration effects across income groups.<sup>6</sup>

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<sup>3</sup>Sommer and Sullivan (2018) study the response of residential investment to nationwide tax policy changes in a full-scale calibrated model with no spatial dimension.

<sup>4</sup>A recent body of work studies the Opportunity Zones created by the Tax Cuts and Jobs Act of 2017 in the United States (Freedman, Khanna, and Neumark, 2021; Chen, Glaeser, and Wessel, 2019; Kennedy and Wheeler, 2021), mostly finding null or modest effects of this large-scale place-based policy on local outcomes, with the exception of Arefeva et al. (2021).

<sup>5</sup>Since Poterba (1984)'s seminal contribution, other papers have explored tax subsidies to home-ownership, from mortgage interest deductions (Hilber and Turner, 2014; Gruber, Jensen, and Kleven, 2021) to home-buyer tax credits (Berger, Turner, and Zwick, 2020; Carozzi, Hilber, and Yu, 2019) or stamp duty exemptions (Best and Kleven, 2018). Studies of the rental market have been mostly confined to tenant-side support for housing consumers, evaluating vouchers (Eriksen and Ross, 2015; Collinson and Ganong, 2018) or cash allowances (Brewer et al., 2019; Fack, 2006). Olsen and Zabel (2015) reviews the literature on housing market policies; see Metcalf (2018) for a survey focused on housing affordability.

<sup>6</sup>Early cross-sectional studies of the LIHTC found close to full crowding-out effects on un-subsidized housing (Malpezzi and Vandell, 2002; Sinai and Waldfoegel, 2005; Eriksen and Rosenthal, 2010). Because the *Pinel* law is targeted at individual – rather than corporate – investors, it has direct implications for the respective role of returns, tax incentives, and spatial frictions in individual landlord decisions. Furthermore, its effects on the local and national housing markets are likely to be distinct from the LIHTC: sales claiming the *Pinel* rebate represented close to half of overall sales of new multi-family units in France in 2016-2018; broad swaths of the population are eligible as tenants; and the mechanism is “open-ended”, while the LIHTC relies on fixed allotments and administrative approvals. Empirically, novel administrative data allow me to examine the role of home-bias at the level of individual units, landlords, and local communities. Finally, while studies of the LIHTC generally rely on *ex post* variation in distance to, and timing of, actual projects implemented, my strategy instead exploits *pre-determined* spatial targeting and sharp eligibility changes over time, without assuming “ring”-based treatment

## 2 Home-biased landlords: three motivating facts

As of 2019, forty percent of households in France rented their homes, with close to two thirds of these tenants residing in about seven million privately provided rental units.<sup>7</sup> Leveraging register data on all dwelling units in France, exhaustive transaction deeds for the entire country, as well as individual asset holdings and income tax returns microdata, this section documents the financial and geographic frictions constraining residential investment by individual landlords. First, most private rental dwellings are leased at low scale by affluent but undiversified “mom-and-pop” investors. Second, properties in the data exhibit highly dispersed rental yields, both across towns and within the same location, with lower rent-to-price ratios in locations located closer to potential investors. Third, landlords display a local bias for nearby properties, with rental ownership probability fast decaying in the distance between investors and their assets. These findings motivate a framework where the location of savers, and their investment behavior sensitivity to returns and spatial frictions, take center stage in the allocation of dwellings and people across space.

**Fact 1: Residential investment is lumpy and operated at low scale by undiversified households** I first use household wealth survey data to document the lumpy, concentrated, and undiversified nature of rental housing ownership.<sup>8</sup> Linking asset ownership information to landlord status evidences several relevant features of the French rental property market. First, as evidenced in panel (a) of figure 1, most landlords in France are so-called “mom-and-pop” investors, who own two or fewer housing units. Rental property is highly lumpy and frictional: two thirds of investors own only one unit, and, as panel (b) of figure 1 illustrates, close to seventy-five percent of all dwellings are owned by landlords possessing two units or fewer.<sup>9</sup> Second, the close to 3 million landlord households are substantially wealthier than both renters and non-landlord owner-occupiers,<sup>10</sup> as shown in panel (a) of figure C.9. The median wealth of landlords is above the 99<sup>th</sup> percentile of net worth for renters: this entails a clear wealth segmentation between the population of moderate-income renters and the more well-off group of owners of the housing stock.<sup>11</sup> Finally, panel (b) of figure C.9 computes

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intensity decay. This paper also complements evidence on the inclusionary zoning participation of developers (Soltas, 2020), but focuses on the intensive margin of housing production, and the allocation of their residential asset portfolio by individual savers.

<sup>7</sup>Aggregate figures are computed from the *Base Logements*, a combination of the Housing survey and Census conducted by France’s national statistical agency (INSEE). The remaining third of tenants live in publicly provided “social” housing, offered by the national governments at deep discounts against market rents. In the United States, c. 36 percent of households were renters as of 2019Q4, according to Census data.

<sup>8</sup>Appendix A provides additional information on the restricted-access microdata from the 2018 household wealth survey (*Enquete Patrimoine*) used in this subsection.

<sup>9</sup>The implied total number of units leased by households in the wealth survey (4.3 million units) is not equal to the total number of rental dwellings owned by private investors in the individual dwellings data (around 7 million in metropolitan France as of 2018). The discrepancy is likely explained by three main factors: non-response, lack of representativeness, or measurement error in the survey data (e.g. when investors state ownership of one dwelling but actually mean a multi-family building with several units); ownership by institutional investors, which is limited and represents between one and two percent of the overall private residential rental housing stock in France, according to INSEE; and foreign ownership of rental housing. However, since the wealth survey data are mostly used to document investor characteristics, rather than properties of the overall housing market, this partial lack of coverage is not consequential for the regularities evidenced here.

<sup>10</sup>An overwhelming majority of landlords in the data (close to 90 percent) are themselves owner-occupiers.

<sup>11</sup>The median net worth of landlords is about €500,000 in 2018, while the corresponding figure for non-landlords owner-occupiers is €217,000, and €10,000 only for renters. Appendix C.2 provides some additional demographic information on the characteristics of owners

the portfolio share of wealth in rental dwellings among landlords. The median share is 52 percent, and rental assets exceed 30 percent of wealth for three quarters of owners. For most landlords, the value of rental property constitutes a substantial proportion of their assets, indicating indivisibilities and frictions to risk-sharing and portfolio diversification. Overall, similar to the way financial frictions can lead to a misallocation of capital in the production of tradable goods (Aghion and Bolton, 1997; Moll, 2014), operational barriers prevent rental housing from being operated at large scale by the most efficient landlords. The lumpy, undiversified, and individual nature of rental ownership makes the spatial distribution of ownership relevant to aggregate rental productivity and housing affordability.

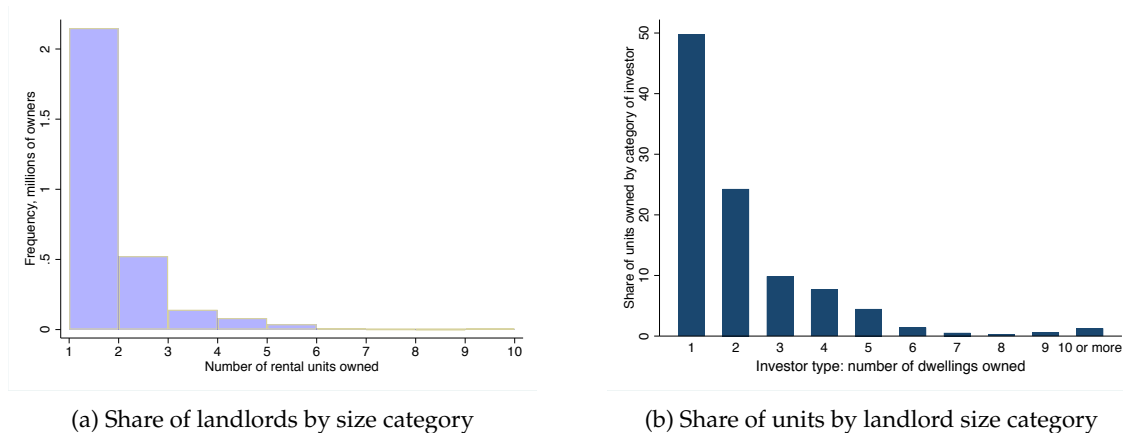


Figure 1: **Landlords and the lumpy ownership of rental housing** Panel a plots the distribution of landlords according to the number of rental units they state to own, using matched household-assets data from the wealth survey. Panel b plots the distribution of units owned by landlord size category. Data are computed at the household level, from the *Enquete Patrimoine* 2017-2018.

**Fact 2: Rates of return are unequal across space and units, and decrease with proximity to investors** I next quantify the large dispersion in contemporary capitalization rates (the ratio of current rents to housing prices) across local housing markets, as well as within them, to provide further evidence of frictions in the allocation of residential investment. I first use town-level 2018 data on rents<sup>12</sup>, and calculate the median acquisition price per square meter for multi-family units in exhaustive transactions deeds data for the three years immediately surrounding 2018, to compute rent-price yields at the municipality level.<sup>13</sup> Although rents and sales prices are highly positively correlated across cities, rent-price ratios still vary widely across space, as shown in panel (a) of figure 2. Capitalization rates are not scattered at random across locations; instead, as shown in panel (a) of figure E.2, I find systematically lower rent-price ratios in cities with higher prices.

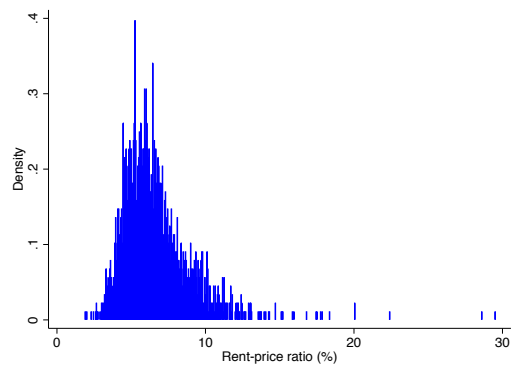
Gross yield dispersion across municipalities, however, may be driven by variation in the average vintage of current local leases, or differential expectations of capital gains and rent growth. Moreover, rentals may differ

of rental dwellings, who are older and have higher income than both non-landlords owner-occupiers and renters.

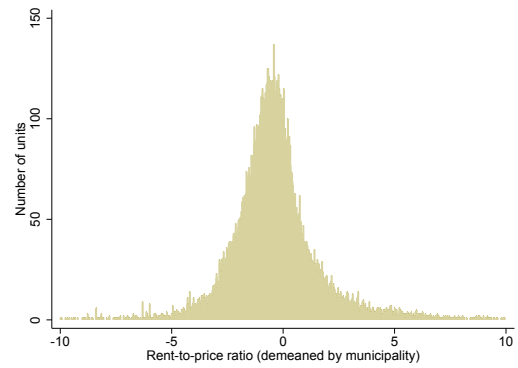
<sup>12</sup>No systematic panel data on rents across locations exists in France. Only one official cross-sectional measure of representative 2018 rents per square meter in multi-family apartments across municipalities, is made available by the French Housing secretariat.

<sup>13</sup>Municipalities, or *communes*, are the smallest French administrative unit, often comparable to US ZIP codes. Their average area is 15 square kilometers (average population of 1,850), although their size varies widely: the largest one, Paris, counts 2.1 million residents. The transactions deeds data are described further in appendix A. To limit measurement error concerns, I focus here on municipalities with more than 11 transactions for multi-family units over the three years period. Appendix figure E.3 verifies that the results are not driven by the selection of larger municipalities, providing similar evidence for all towns with no restriction on the number of transactions.





(a) Rent-to-price: across municipalities



(b) Rent-to-price, buy-to-let units: within municipalities

**Figure 2: Dispersion in rent-to-price ratios** Figure 2 documents the dispersion in rent-to-price ratios across and within municipalities. Panel a plots the distribution of rent-to-price ratios across municipalities, using average rents per square meter for multi-family units in 2018, and median acquisition prices for multi-family units from 2017 to 2019 in transaction deeds data (*DV3F* database). Panel b plots the distribution of re-centered (by municipality-year level mean) rent-to-price ratios across rental units purchased under the *Pinel* subsidy scheme, for municipalities with more than one unit, using individual tax returns (*POTE* database) for all purchases made in 2016 under the *Pinel* scheme in 2016 (see main text).

from recently transacted units due to selection bias (Eichholtz et al., 2021), making cross-city rent-price-ratio comparisons difficult.<sup>14</sup> To illustrate the large residual variation in rental returns even after controlling for expected local dynamics, I turn to a measure of the dispersion of within-location-year unit-level gross rental yields, using confidential individual tax returns. To do so, I exploit the fact that households who purchase a buy-to-let unit under a subsidized housing scheme in France must fill out a “buyer’s commitment form” stating the purchase price and initial rent of the unit.<sup>15</sup> I compute rent-to-price ratios for close to 65,000 housing units purchased under the *Pinel* subsidy scheme in 2016. Demeaning rent-price ratios by a municipality-year level average, the distribution of re-centered rates of return in panel (b) of figure 2 still exhibits substantial dispersion in returns within a location.<sup>16</sup> While maximum rents per square meter are regulated by a ceiling under the *Pinel* subsidy scheme (see *infra*, section 4), both actual rents per square meter and purchase prices vary widely within towns. Similar to the cross-city evidence, panel (b) of figure E.2 shows rent-to-price ratios are lower in higher-price units, even after accounting for municipality-level fixed effects.<sup>17</sup>

Moreover, beyond the large degree of dispersion, I show that the rent-price ratio in a location is negatively correlated with the proximity of the city to potential investors. Using income tax data aggregated across France’s

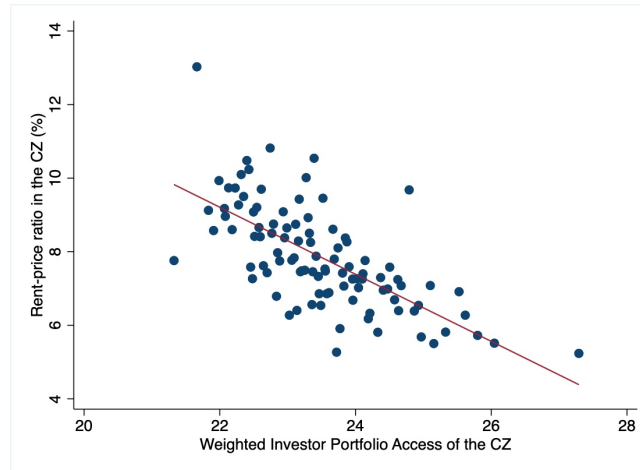
<sup>14</sup>In a standard Campbell-Shiller decomposition (Campbell and Shiller, 1988) decomposition, current rent-price ratios reflect expected changes in future rents and capital gains stemming from changes in discount rates (Campbell et al., 2009). Within-lease rent changes are heavily regulated in France. For the same tenant, the maximum annual increase is limited to the yearly evolution of a nationwide official index (the “*Indice de Reference des Loyers*” or *IRL*) computed on the basis of the overall CPI (excl. rents and tobacco). In case of a tenant change, the rent can increase by more than the *IRL*, except since 2012 in so-called “tense” property markets covering large cities, where new-lease rent increases are also limited by the *IRL*, unless the property has been vacant for more than eighteen months or has undergone substantial improvement works. As regards capital gains, a standard model with a representative owner and mean-reverting prices suggests that higher future capital gains should be correlated with lower current prices, leading to counter-factually higher rent-price ratios in low price locations (Sivitanides et al., 2001; Liu, Wang, and Zha, 2019).

<sup>15</sup>Buyers must fill this form, known as *schedule 2044EB*, when they first file their individual income tax returns in the year of the purchase, in order to benefit from the conditional rental investment tax credit. The form includes the purchase price of the unit, its location, the purchase date, as well as the initial rent at which it will be leased to the tenant.

<sup>16</sup>City fixed-effects only account for 13 percent of the overall variance in rent-price ratios in the data consisting of buyers’ commitment forms in individual tax returns.

<sup>17</sup>Asset-level variation in gross rental yields even within narrow locations accords with the idiosyncratic risk in Giacoletti (2021).

close to 300 commuting zones, I first compute a proxy for local wealth  $W_i$  in location  $i$  as the sum of local capital income earned by residents of  $i$ ,  $W_i \propto \tilde{W}_i = \sum_{z \in i} y_z^K$ .<sup>18</sup> I then calculate an empirical proxy for the “investor portfolio access” of location  $j$ , an index for its physical proximity to investors, as the sum of wealth in all locations  $i$ , inverse-weighted by a function of their bilateral distance to  $j$ ,  $d_{ij}$ :  $IPA_j = \sum_i e^{-d_{ij}} W_i$ .<sup>19</sup> Figure 3 evidences that a commuting zone’s rent-to-price ratio systematically and negatively correlates with investor proximity: cities that are physically closer to savers exhibit lower equilibrium rent-price ratios.



**Figure 3: Rent-price ratios and investor proximity** Figure 3 plots a binned scatter plot of the relationship between the rent-to-price ratio,  $\frac{R_j}{P_j}$ , and (the log of) the estimated “investor portfolio access”  $IPA_j = \sum_i W_i e^{-d_{ij}}$  on the x-axis, where  $d_{ij}$  is the bilateral haversine distance between CZs  $i$  and  $j$ , and  $W_i$  is a measure of total capital income in  $i$ . The figure is estimated across France’s 293 mainland commuting zones with available data and town-level rent-price ratios within a CZ are aggregated after weighting by the number of transactions over 2017-2019.

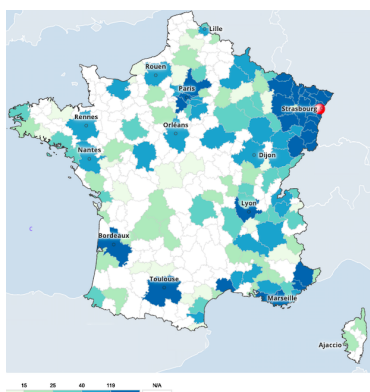
Taken at face value to embody variation in real returns to housing investment,<sup>20</sup> the dispersion in rent-price ratios itself (figure 2) suggests a “Lucas paradox” of residential investment: why does it not flow from low to high rate-of-return locations and units? While merely suggestive of the presence of distance frictions, the decreasing relationship of rent-price ratios with proximity to investors (figure 3) offers one potential explanation, further explored in my stylized model: the spatial segmentation of rental investment may entail an inelastic supply of imperfectly mobile capital, which fails to equalize returns across space.

<sup>18</sup>Local wealth data are not directly available, because the wealth survey does not record precise household location information. I use aggregate taxable capital income from administrative data at the commuting zone level, as a proxy for the ranking of locations by wealth. In the spirit of the wealth inequality literature’s common use of the “income capitalization” method (King, 1927; Stewart, 1939), under the extreme first-order approximation of a single capitalization factor, log wealth in a location would be directly proportional to log capital income in a location (Chodorow-Reich, Nenov, and Simsek, 2021).

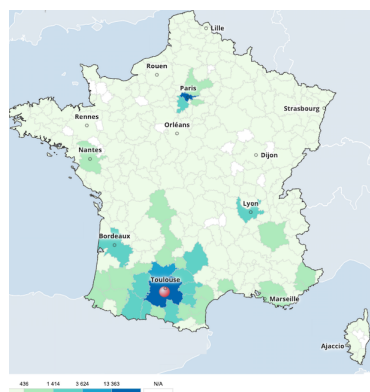
<sup>19</sup>The choice of aggregation at the CZ level relates to the computational feasibility of measuring all possible distances across location pairs: for close to 300 mainland CZs, the number of bilateral pairwise distances to estimate is about 45,000, but it would rise exponentially when estimating the entire network of pairwise distances across France’s 35,000 municipalities. Anticipating on the predictions of the framework in section 3, I take an inverse exponential function as my baseline approach to inverse-distance weighting of access to wealth.

<sup>20</sup>As in the international finance literature on investment frictions (Caselli and Feyrer, 2007), observed returns may be subject to measurement error, which can shroud the true variation across locations and units. I take the substantial observed dispersion in rental rates, and their systematic decreasing pattern with investor portfolio access, to be a stylized regularity consistent with impediments to the flow of capital across cities in the rental market.

**Fact 3: There is a “home bias in homes” and landlord investment decays with distance** Lastly, I document spatial constraints to rental property investment directly from exhaustive register data on the universe of rental properties in France and their owners. In the absence of any distance barriers in the rental market, the residence of landlords owning property in any given city would follow the same spatial pattern as the overall landlord population, as a byproduct of portfolio diversification. In practice, however, this null hypothesis of perfect diversification is strongly rejected. Housing investment exhibits substantial home bias towards areas close to the residence location of the landlord-investor. The count of units  $C_{ij}$  in rental market  $j$  owned by investors living in  $i$  decays substantially with bilateral geographic and social distance  $d_{ij}$  between the two cities.<sup>21</sup>



(a) Ownership of Strasbourg rental units



(b) Ownership of Toulouse rental units

**Figure 4: The spatial pattern of rental asset ownership** Figure 4, panel a (resp. b), plots, for each of 304 commuting zones in mainland France and Corsica, the number of units owned in CZ  $j$  : Strasbourg (resp.  $j$  : Toulouse) by landlords residing in any CZ  $i$  ( $C_{ij}$ ), including in the same commuting zone. CZs are ordered into 5 classes of ownership intensity of destination CZ units by origin CZ investors, shaded from lightest to darkest, according to the Jenks-Fisher natural breaks optimization routine. In accordance with data disclosure agreements with the French Treasury, only CZs where landlords own more than 11 units in the destination rental market are shown. Statistics are computed from individual ownership data for the entire stock of privately owned rental units in France (the 2019 FIDELI database for c. 7 million rental dwellings used as primary residences).

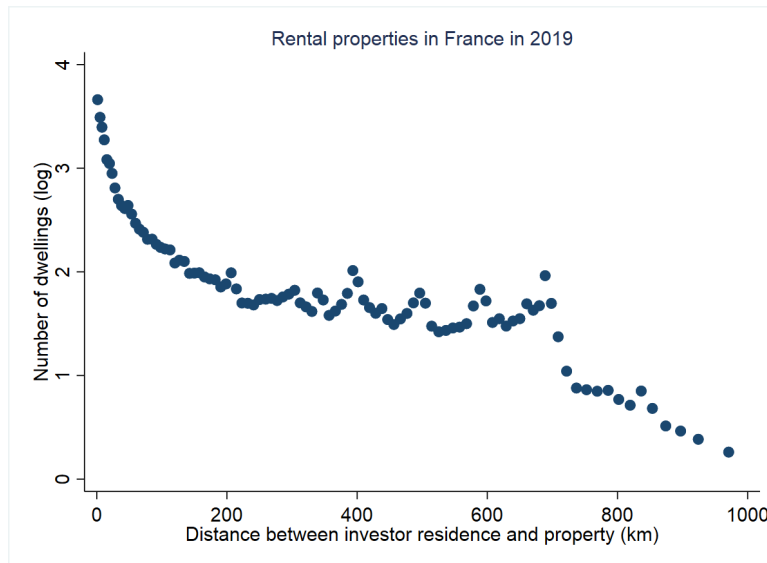
Panels (a) and (b) of figure 4 illustrate this fact, mapping the residence distribution of individual owners of rental dwellings located in two large French commuting zones, Strasbourg and Toulouse.<sup>22</sup> While owners from populated and wealthy commuting zones like Paris or Lyon own a large number of units in both destinations, the striking pattern is the presence of substantial local bias, with the share of owners living in the commuting zone itself above 60 percent, and a clear geographic decay of portfolio shares with the distance between investor residence  $i$  and destination  $j$ . This more systematic finding is documented for close to 7 million private rentals in France in figure 5: the (log) number of units owned by investors from  $i$  in  $j$  decreases strongly with the distance between  $i$  and  $j$ .<sup>23</sup> Table 1 presents “gravity” estimates of the distance effect on rental

<sup>21</sup>This regularity is reminiscent of the international finance literature, where foreign direct investment and portfolio holdings have been found to increase with measures of bilateral country proximity (Portes, Rey, and Oh, 2001; Portes and Rey, 2005). The presence of within-country home bias in investment also accords with the findings of Coval and Moskowitz (1999) for institutional investors stockholdings.

<sup>22</sup>The number of rentals owned in  $j$  for owners residing in  $i$  is measured from exhaustive cadaster information in the FIDELI database (see appendix A), aggregated at the level of mainland France’s 304 commuting zones (“zones d’emploi”).

<sup>23</sup>I compute the haversine distance between the town in which a property is located, and the town of the address to which the property tax is sent (a proxy for the residence of the owner, since owners are legally liable for property tax payments in rental units), for all private rental units in the country.

property ownership, conditional on origin and destination fixed effects. The distance coefficient for predicting relative ownership probabilities hovers around unity; it is significant and economically substantial, with the effect largely mediated by social proximity. Figure i visualizes home bias for alternative distance measures (haversine distance, bilateral road distance) and aggregation levels (provinces, CZ).<sup>24</sup>



**Figure 5: Distance effects on ownership of rental housing** Figure 5 documents the spatial concentration of housing ownership, using individual ownership data for the entire stock of privately owned rental units in France. It plots a binned scatter plot of the (log) number of units in municipality  $j$  owned by households residing in a municipality  $i$ , against bins of geographic haversine distance (in km, plotted up to a maximum distance of 1,000km) between the centroids of municipalities  $i$  and  $j$ . The municipality of residence of the owner is imputed from the municipality of the address to which the property tax (remitted by the owner) is sent.

The over-representation of local ownership in the rental market is even more puzzling than international equity home bias. Most landlords are owner-occupiers themselves: they could better insure against local price risk by investing in distant locations with low correlation to domestic house prices.<sup>25</sup> Several factors could account for the strong observed home bias. Local landlords may have informational advantages relative to out-of-town buyers, through better timing, targeting, or negotiating strategies (Chinco and Mayer, 2016); they could save on intermediation and maintenance costs; or they could be better able to monitor tenants and alleviate the moral hazard inherent to the provision of rentals.<sup>26</sup> Without taking a stand on the underlying explanation, the large effect of distance on ownership suggests an imperfect mobility of rental capital supply across space: frictions reduce the subjective returns of landlords when investing in homes away from home.

<sup>24</sup>Social connections are proxied by the Facebook friendship-based Social Connectedness Index (Bailey et al., 2018), only available across NUTS2-level provinces (*departements*) in France.

<sup>25</sup>Home-ownership itself may provide a long-term hedge against volatile local rents, for owners endowed with location-specific human capital (Sinai and Souleles, 2005). However, conditional on hedging rent risk in their own preferred location via owner-occupied housing, the ownership of *additional* rental assets in or near the same city is less straightforward to rationalize under perfect capital mobility.

<sup>26</sup>An alternative explanation involves long-term persistence. If movers lease, rather than sell, their past owner-occupied homes, and if mobility is limited across space, then the ownership structure of rental assets would exhibit local bias. Similarly, with bequests and low inter-generational spatial mobility, landlords will tend to live close to their parents' former homes, which they inherited and may now lease; and if landlords invest in local assets with a view to donating them to their own children, limited inter-generational spatial mobility and altruistic motives will again entail negative distance effects on rental property ownership. However, for landlords to prefer continued ownership of such rental assets, they must still exhibit some objective or subjective comparative advantage in their operation.

### 3 Home-biased landlords: a spatial model of rental housing

The previous section provided stylized evidence that landlords, who are older and wealthier than the rest of the population, are mostly undiversified individual operators of rentals close to their own residence, facing significant capital market frictions. I outline here how an illustrative conceptual framework combining a standard location choice with the frictional portfolio decisions of home-biased landlords can rationalize the steady-state regularities documented in section 2, and deliver specific predictions for the quasi-experiment analyzed in sections 4 and 5 of this paper.

#### 3.1 Setup: renters, savers, and developers

**Environment** Locations are indexed by  $j = 1, \dots, J$ . A number  $\bar{L}$  of agents are born at time  $t$ , and live for two periods in overlapping generations. Each generation is split into two fixed-types subsets: mobile hand-to-mouth worker-renters  $W$ , and immobile owners-savers  $O$  who can become landlords. Wages  $y_j^k$  (for  $k \in O, W$ <sup>27</sup>) and amenities  $B_j$  in each location are exogenous, while home prices, rents, and population are the endogenous outcome of mobility decisions, housing production, and portfolio allocation choices.

**Workers-renters** A share  $\lambda$  of households rent housing and work in final good production. Worker-renter behavior follows standard spatial equilibrium models (Redding and Rossi-Hansberg, 2017). These agents, indexed by  $\omega$ , choose where to live at birth for the duration of their existence, earning the city-type specific wage income each period. Their utility is given by composite consumption in each of the two periods of their lives:

$$U_j^W(\omega) = (1 - \beta^R) \ln(C_{t,j}^W \eta_j(\omega)) + \beta^R \ln(C_{t+1,j}^W \eta_j(\omega))$$

Renters are hand-to-mouth: they consume their entire wage income each period, and, with Cobb-Douglas preferences over housing and the (numeraire) final good, they spend a share  $\gamma$  of their income on rent. With constant local wages, consumption is the same each period and lifetime utility equals  $\ln(C_{t,j}^W \eta_j(\omega))$ . Workers elect a preferred location depending on wages for their type  $y_j^W$ , housing rent  $R_j$ , and idiosyncratic preference shocks  $\eta_j(\omega)$ , drawn from an extreme-value distribution with shape parameter  $\nu$  and scale  $B_j$  denoting average amenities, as in Redding (2016). Standard arguments imply that the share of renters choosing to reside in  $j$  is given by:

$$\pi_j^W = \frac{B_j (y_j^W R_j^{-\gamma})^\nu}{\sum_k B_k (y_k^W R_k^{-\gamma})^\nu} \quad (1)$$

In turn, the quantity of rental housing services required by a generation of workers-renters in location  $j$  is:<sup>28</sup>

$$H_j^{WR} = \frac{\pi_j^W \times \lambda \bar{L} \times \gamma y_j^W}{R_j} \quad (2)$$

<sup>27</sup>Production in the model is linear in each type of local labor, so that type-specific earnings are constant.

<sup>28</sup>Because I focus on steady-state implications of the model, I mostly omit time subscripts where it is innocuous.

**Landlords-savers** The remaining  $(1 - \lambda)\bar{L}$  agents are fixed-location owners-occupiers. An exogenous share  $\bar{\pi}_i^O = \frac{L_i^O}{(1-\lambda)\bar{L}} > 0$  of them lives in each location  $i$ . To account for their fixed location, one may assume these agents are endowed at birth with unalienable owner-occupied housing in a given city - for example, through bequests from the last generation of owners. Unlike workers-renters, they can – and, in equilibrium, must – save. An owner-saver residing in  $i$ , indexed by  $z$ , earns a city-type-specific income  $y_i^O$  in period  $t$  (“working life”), and lives on the proceeds of his investments in  $t + 1$  (“retirement”). The utility of savers is increasing in consumption in both periods, with discount factor  $\beta_i^O$ :

$$U_i^O(z) = (1 - \beta_i^O) \log(C_{t,i}^O(z)) + \beta_i^O \log(C_{t+1,i}^O(z))$$

Log utility implies each saves a constant fraction  $\beta_i^O$  of their first-period income  $y_i^O$ .<sup>29</sup> During their working life, savers face two potential subsets of assets. First, they can invest in “stocks” (representing the global financial market), receiving a baseline gross return  $r_K$ . Alternatively, they can purchase housing units in destination city  $j$  before leasing them to next-period renters. If they choose to do so, they pay the unit price  $P_j^H$ , before receiving (potentially subsidized) rents of  $R_j S_j$  in the next period,<sup>30</sup> yielding a return  $r_j^H = \frac{R_j S_j}{P_j^H}$ . Each constrained investor allocates their savings to their highest-return investment opportunity.<sup>31</sup> They are, however, subject to frictions, so that individual subjective returns from an investment opportunity differ from market-level observed gross return rates.<sup>32</sup> Closely following Pellegrino, Spolaore, and Wacziarg (2021), the subjective returns to an investor  $z$ , living in  $i$ , from investing in an asset  $\ell$  (a rental dwelling in location  $j$  delivering  $r^\ell = r_j^H$ , or a stock with  $r^\ell = r_K$ ), are altered by an “iceberg” friction:

$$r_{z,\ell}(i, j) = r^\ell e^{-\lambda(z,\ell,i,j)}$$

where:

$$\lambda(z, \ell, i, j) = \delta \times \mathbb{1}(\ell \in H) \times d_{ij} - \epsilon_{z,i}$$

<sup>29</sup>In the absence of risk, this functional form choice for period felicity delivers constant savings shares independently of local expected returns, since income and substitution effects of varying asset returns compensate each other exactly.

<sup>30</sup>In the natural experiment of in the next sections, investment tax credits in  $j$  lower the purchase price for investors by  $(1 - s_j)$ , but require rents to remain below a ceiling  $\bar{R}_j$ . The combination of these tools (a subsidy to the purchase price and a reduction in rents) in this simple model is isomorphic to a proportional subsidy  $S_j$  on rents received by landlords. Subsidies are funded with lump-sum nationwide taxes that do not distort worker location or saver investment decisions.

<sup>31</sup>To make the implications of the framework stark, and consistent with motivating facts on landlord wealth documented earlier, I assume savers are unable to borrow when taking advantage of high-return opportunities, so that their investment decision is constrained by their own savings – but I conjecture that any binding contractual frictions limiting their ability to access external capital to a multiple of their own private savings (as would result from standard repayment monitoring constraints a la Holmstrom and Tirole (1997)) would have similar qualitative consequences.

<sup>32</sup>In practice, in the case of the housing market, subjective return variation from a housing investment project in a given destination city may correspond to any investor- and asset-specific idiosyncratic payoffs. As mentioned in section 2, these could stand for an individually-varying hassle cost to examine, purchase or maintain the property. Alternatively, akin to the differentiated liquidity benefits assumed in the literature on “convenience yields” for financial securities (Feenstra, 1986; Krishnamurthy and Vissing-Jorgensen, 2012), they could represent individual variation in the perceived probability of enjoying a given asset as a secondary residence, or of bequeathing it to one’s children. For other outside assets (“stocks”), any variation in individual-asset specific transaction costs or individual preferences could generate the type of return dispersion assumed here. Absent any financial frictions, only investors with the highest subjective return would invest in a given asset, borrowing from all other savers, an implication that runs counter to observed patterns of portfolio allocation.

In the housing sector, investment frictions involve a systematic component increasing with the bilateral distance  $d_{ij}$  between the origin and destination locations. Moreover, all returns are affected by an idiosyncratic shock  $\epsilon(z, \ell)$ , at the saver  $z$ -asset  $\ell$  pair level, drawn from a nested Fréchet distribution.<sup>33</sup> Since each saver allocates their entire savings to their highest return opportunity, this delivers a tractable expression for  $w_{ij}$ , the share of investors from  $i$  who purchase residential assets in  $j$ . This proportion is also, by the law of large numbers, the share of wealth in  $i$  ( $W_i = (1 - \lambda)\bar{L}\pi_i^O\beta_i^O y_i^O$ ) invested in  $j$  housing:

$$w_{ij} = \underbrace{\frac{\tilde{m}_{Hi}^\phi}{\tilde{m}_{Hi}^\phi + r_K^\phi}}_{\text{Housing share } HS_i \text{ in } i \text{ assets}} \times \underbrace{\frac{\left(e^{-\delta d_{ij}} r_j^H\right)^\theta}{\sum_{j'} \left(e^{-\delta d_{ij'}} r_{j'}^H\right)^\theta}}_{\text{Share } s_{ij} \text{ of } j \text{ in } i\text{'s housing portfolio}} \quad (3)$$

where  $\tilde{m}_{Hi} = \left(\sum_{j'} \left(e^{-\delta d_{ij'}} r_{j'}^H\right)^\theta\right)^{\frac{1}{\theta}}$  is a theory-consistent index of net housing returns for investors living in  $i$ . The portfolio weight  $w_{ij}$  is the product of two terms. The first is the overall probability  $HS_i$  of investing in housing (rather than stocks) for investors from  $i$ ; the second is the probability (for  $i$  savers) of buying property in  $j$ , conditional on investing in any housing asset. A higher  $\theta$  makes housing investment more elastic to returns across space, while a higher  $\phi$  makes housing and non-housing investments more substitutable from the point of view of investors. The demand for housing assets in location  $j$  at time  $t$  (which is the time- $t + 1$  supply of rental services) is given by the aggregation of investors' purchases from each location  $i$  ( $H_{i \rightarrow j}$ ):

$$P_{jt}^H H_{jt} = P_{jt}^H H_{j,t+1}^{SR} = \sum_i P_j^H H_{i \rightarrow j} = \sum_i \pi_i^O (1 - \lambda) \bar{L} \times \beta_i^O y_i^O \times w_{ij} \quad (4)$$

**Housing production by developers** Housing used for rental services is produced by competitive developers using local land  $T_j$ , as well as materials  $X_j$  (units of the numeraire). The price of local land is  $r_j^T$ , with a limited supply assumed to depend on exogenous characteristics  $\bar{T}_j$  and its own price:  $T_j = \bar{T}_j (r_j^T)^{\sigma_j}$ . The production function for new housing is Cobb-Douglas, with land share  $\xi_j$ :  $H_j = \left(\frac{T_j}{\xi_j}\right)^{\xi_j} \left(\frac{X_j}{1 - \xi_j}\right)^{1 - \xi_j}$ . Solving for the maximization problem of competitive developers, the land price is given by  $r_j^T = (P_j^H)^{\frac{1}{\xi_j}}$ , and the supply of housing assets can be shown to follow<sup>34</sup>:

$$H_j = \frac{\bar{T}_j}{\xi_j} (P_j^H)^{\frac{1 + \sigma_j - \xi_j}{\xi_j}} \quad (5)$$

<sup>33</sup>Denoting each sector (housing and outside assets) by  $s$ , and  $N_s$  the number of potential assets in each sector, the distribution of shocks satisfies:

$$F(\epsilon_{z,t}(s, k)) = \exp \left[ - \sum_{s \in (K, H)} \left( \sum_{k=1}^{N_s} (\epsilon_{z,t}(s, k))^{-\theta} \right)^{\frac{\phi}{\theta}} \right]$$

This type of nested shock distribution, with individual extreme-value correlated preference shocks giving rise to a nested logit choice probability (Verboven, 1996), has been used to model preferences across goods and varieties (Atkeson and Burstein, 2008), firms' funding choices (Herreno, 2020), or workers' employment location decisions (Berger, Herkenhoff, and Mongey, forthcoming). The (inverse) dispersion of shocks across the housing and non-housing nests is parametrized by  $\phi$ ; while the (inverse) dispersion of return shocks within housing assets depends on  $\theta$ .

<sup>34</sup>Equilibrium land prices are log-linear in local housing prices since materials are the numeraire; and the output price elasticity of housing production is increasing in the land conversion responsiveness, and decreasing in the land share of construction.

**Definition: stationary equilibrium** Given wages  $y_j^W, y_j^O$ , amenities  $B_j$ , buildable land characteristics  $\bar{T}_j$ , subsidies  $S_j$ , and owner residence distribution  $\pi_j^O$ , a steady-state spatial equilibrium of this economy with constant population is an allocation of renters  $L_j^W$ , housing production  $H_j$ , housing prices  $P_j^H$ , and rents  $R_j$  such that, in each generation:

- The number of workers-renters living in each destination market  $j$ ,  $L_j^W$ , is constant and satisfies their optimal location choice, summarized by  $\pi_j^W$  in equation 1.
- The savings allocated by investors from each origin city to each destination housing market are constant, and, given returns, satisfy the optimality condition for individual investors, summarized by the weights  $w_{ij}$  in equation 3.
- The demand for new housing assets at time  $t$  is given by investor asset allocation 4; the supply of new housing is determined by developer competition 5; and the market for new housing production and land both clear each period.
- The market for rental services clears each period, with the demand  $H_{j,t}^{WR}$  given by renter aggregation 2 at time  $t$ , and the supply  $H_{j,t}^{SR}$  given by portfolio allocation choices 4 at time  $t - 1$ .

**Discussion of assumptions** The framework outlined above is highly stylized, but the hypotheses made on renters and landlords are rooted in real-world regularities.

First, the model postulates that only renters move across space. Using Census data, appendix C.1 provides a trove of empirical evidence at the individual and town level that renters are much more prone to moving than owner-occupiers, likely due to the presence of transaction and mobility costs. Tenure duration in the private rental sector is substantially lower than in both public rental and owner-occupied housing segments.

Second, the model assumes a perfect segmentation of the population between fixed-location, "older", landlords, and mobile, hand-to-mouth, "younger" renters. While clearly an extreme case, since many households may transition from renting to owning over the course of their life cycle, this distinction is qualitatively consistent with the descriptive stylized facts presented in section 2 and table C.4 in appendix C.2: individual landlords are substantially older and wealthier than renters – whose savings are almost nil.

Third, the framework distinguishes between the production of housing by developers, and the delivery of rental services by landlords. This structure is peculiarly suited to the analysis of the French (and Western European) rental market, where builders sell ownership rights for dwellings to individual investors before or immediately after construction. The United States has a larger proportion of corporate landlords who erect, own and operate entire multi-family buildings, but individual landlords still play a substantial role, representing 7 percent of all US households and operating 41 percent of all rental units.<sup>35</sup> The separation of competitive production, on the one hand, and lumpy ownership and operation, on the other, is also relevant in the US.

Fourth, instead of a standard portfolio choice under risk, spatial frictions to the operation of housing assets are assumed to reduce the private value of operating more distant rental investments, consistent with the home bias documented in section 2. This geographic segmentation also implies that many local owners of rental property may be infra-marginal in their portfolio allocation.

<sup>35</sup>These summary data for the rental market in the United States are computed from the Census's Rental Housing Finance Survey for 2018 and the IRS's Statistics of Income for the same year.



### 3.2 Model implications: the steady-state

**The cross-section of cities** The model matches key steady-state regularities of the rental market documented in section 2. By making assumptions on the frictional nature of housing investment across locations, it generates a cross-sectional spatial pattern of ownership and rent-to-price ratios consistent with the data.

**Proposition 3.1.** (i) In equilibrium, the network for housing asset ownership exhibits a gravity structure, summarized by:  $H_{i \rightarrow j} = A_i \times \tilde{A}_j \times e^{-\delta \theta d_{ij}}$ . (ii) In a two-city symmetric version of the model, where two locations A and B only differ only by their share of saver wealth  $\pi_A^O y_A^O > \pi_B^O y_B^O$ , the city with a higher "investor portfolio access" (A) exhibits a lower rent-to-price ratio  $\frac{R_A}{P_A} < \frac{R_B}{P_B}$ , a larger housing supply ( $H_A^W > H_B^W$ ) and a higher population of workers-renters ( $L_A^W > L_B^W$ ). (iii) These cross-sectional differences vanish in the absence of distance frictions ( $\delta \rightarrow 0$ ).

- *Gravity in housing investment* Because of the dispersion of owners' idiosyncratic preferences, and the presence of distance frictions in investment, the model delivers a gravity structure of the rental housing ownership network similar to the one observed in the third fact of section 2 and quantified in table 1. Specifically, housing in  $j$  owned by investors from  $i$  has the following gravity structure<sup>36</sup>:

$$H_{i \rightarrow j} = \underbrace{\kappa \times \pi_i^O (1 - \lambda) \bar{L} \times \beta_i^O y_i^O \frac{\tilde{m}_{Hi}^{\phi - \theta}}{\tilde{m}_{Hi}^\phi + r_K^\phi}}_{\text{Origin FE: } A_i} \times \underbrace{\frac{(H_j)^{-\theta \left( \frac{\xi_j}{1 + \sigma_j - \xi_j} + \frac{1}{1 + \gamma \nu} \right)} (B_j (y_j^W)^{1 + \nu})^{\frac{\theta}{1 + \gamma \nu}}}{\left( \frac{\xi_j}{T_j} \right)^{\frac{\xi_j \theta}{1 + \sigma_j - \xi_j}}}}_{\text{Destination FE: } \tilde{A}_j} \times \underbrace{e^{-\delta \theta d_{ij}}}_{\text{Bilateral friction}}$$

- *Decreasing rent-price ratios with proximity to wealth* Locations in close proximity to savers feature a "savings glut" in the housing sector, with a large supply of capital and low returns.<sup>37</sup> While the proposition is stated for the symmetric case of two cities, the logic is more general. More numerous (high  $\pi_i^O$ ) or wealthier (high  $y_i^O$ ) savers in locations  $i$  "close" to  $j$  (i.e. with low effective bilateral frictions  $d_{ij}$ ) supply relatively more potential housing capital to  $j$ . All else equal, this improved "investor portfolio access" (summarized by a higher  $\sum_i e^{-\delta d_{ij}} \beta_j^O \pi_j^O y_j^O$ ) lowers the equilibrium return to housing in  $j$ .<sup>38</sup>
- *Poor people in wealthy cities* Spatial equilibrium models often ponder how to account for the observed presence of low-income workers in expensive agglomerations, since competition for land may be expected to induce full segregation of types.<sup>39</sup> With segmented capital markets, a larger number of affluent savers directly lowers the rent-price ratio, a pecuniary cross-agglomeration force pushing for the in-migration of low-income renters – conditional on other features of housing production. Locations with a larger supply

<sup>36</sup> $\kappa$  is a composite constant dependent on model parameters.

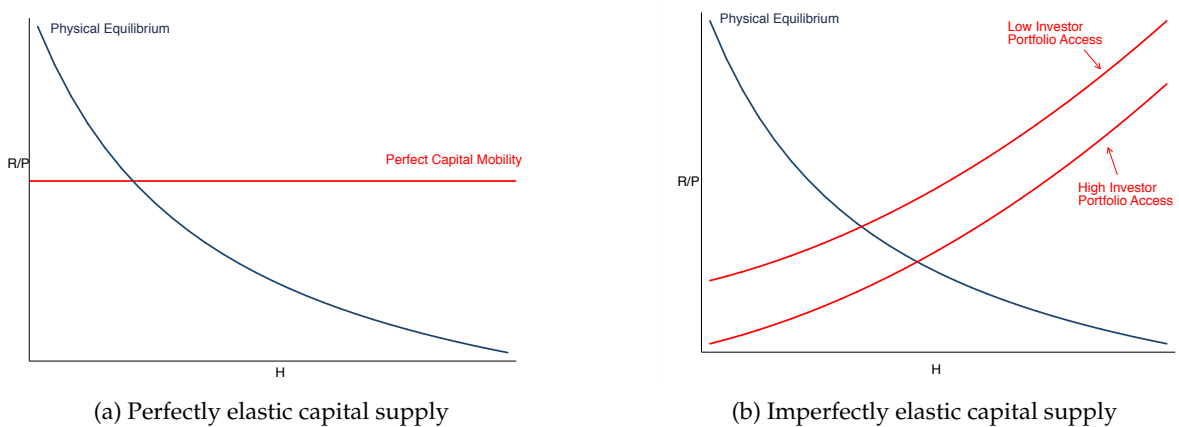
<sup>37</sup>This relates the findings in my paper to recent evidence in Kuchler et al. ([forthcoming](#)) on how higher social connections to locations replete with institutional investors tend to increase firm valuation. This "investor portfolio access" effect is also similar to how, in the presence of trade in goods, locations where transportation infrastructure provides a high market access to efficient producers feature lower effective consumer price indices (Donaldson and Hornbeck, 2016).

<sup>38</sup>Separately, in the many-cities version, the net effect of investor portfolio access on rents and on prices respectively is ambiguous, as it depends on the respective elasticities of land supply  $\sigma_j$  and location choices  $\nu$ , along with the investment reallocation parameters  $\theta$  and  $\phi$ .

<sup>39</sup>Asking this exact question, Glaeser, Kahn, and Rappaport (2008) provide an alternative answer focusing on the role of access to public transportation.

of rental housing (as a consequence of their proximity to investors) host more mobile (lower-income and younger) workers, consistent with the evidence documented in appendix C.1.<sup>40</sup>

**Graphical intuition for proposition 3.1** One can visualize the steady-state of this economy as the combination of a real or “physical” spatial equilibrium schedule, in which workers make location choices as a function of rents and developers supply housing units as a function of prices; and a “financial” optimality condition, in which savers allocate their wealth to housing assets depending on their returns (net of subsidies and frictions).



**Figure 6: Rent-price ratios and rental supply: steady-state** Figure 6 summarizes the role of imperfectly elastic capital supply for the equilibrium supply of rentals (on the x-axis) and rent-price ratio (plotted on the y-axis). In both panels, the decreasing schedule (blue line) summarizes the “physical” equilibrium (a higher quantity of rentals leads to lower rents and higher prices, and thus to a lower rent-price ratio). The capital supply schedule is plotted (in red) for two cases. Panel (a) presents a generic city in a financially integrated country with an exogenous and constant required return for landlords. Panel (b) plots the case of an imperfectly elastic supply of capital, and shows the impact of spatial frictions on steady-state returns and supplies of rentals for two locations with (resp.) a high (low) equilibrium access to investors’ portfolios.

On the “physical” side, all else equal, a larger quantity of rental housing in a location requires *lower* rents to clear the market for rental services (“going down” the demand curve of workers for rentals), and *higher* unit prices to clear the market for housing production (“going up” the supply curve of developers for new units). Combining these, the “physical” equilibrium delineates a *decreasing* schedule between local housing quantities and the rent-to-price ratio.

In the absence of financial frictions, the intersection of this schedule with an exogenously defined required return  $r^K$  for absentee landlords would equalize the rent-to-price ratio in every location. With imperfect capital mobility, however, the “financial” side delivers a location-specific upwards-sloping supply of rental capital in the local rent-price ratio: attracting further investment to a city (by drawing in more distant investors and reallocating non-housing capital) requires higher gross returns. As a consequence of this origin-destination-

<sup>40</sup>In an extended version of the model with mobile landlords, a dynamic counterpart of this prediction would exist. If landlords were to also move in the presence of amenity differentials, their share in  $j$ ,  $\pi_j^O$ , would rise in response to an increase in local amenities  $B_j$ . Assuming such an amenity shock initially draws both landlords and workers to  $j$ , savers now face lower frictions to rental investment in  $j$  on average. This pushes the rent-to-price ratio lower in  $j$  (and higher in locations experiencing reduced investor portfolio access), dampening the direct impact of the population inflow on rental costs in the city. This force may partially account for the long-run drop in rent-price ratios specific to “superstar cities”, where prices have risen faster than rents (Gyourko, Mayer, and Sinai, 2013; Hilber, Mense, et al., 2021).

specific capital supply, locations with a higher aggregated “investor portfolio access”, a measure of their proximity to investors, feature a capital supply curve shifted outwards, and lower returns to rental investment. Figure 6 summarizes this discussion: the left panel describes the case of integrated capital markets for a generic city, while the right panel shows local quantities and rent-price ratios for two cities, with a low or high access to investors’ portfolios, under imperfect capital mobility.

### 3.3 Model implications: the effect of subsidies

**Response to landlord subsidies** The model generates direct predictions for the consequences of raising  $S_j$ , the subsidy to rental investment in location  $j$ , to which I turn in the quasi-experimental evidence obtained in sections 4 and 5. In particular, the theory predicts not only that subsidies raise investment and housing supply in the targeted location, but also that their effects operate through a rise in long-distance investor involvement, and are stronger in locations exhibiting more openness to outside investors.

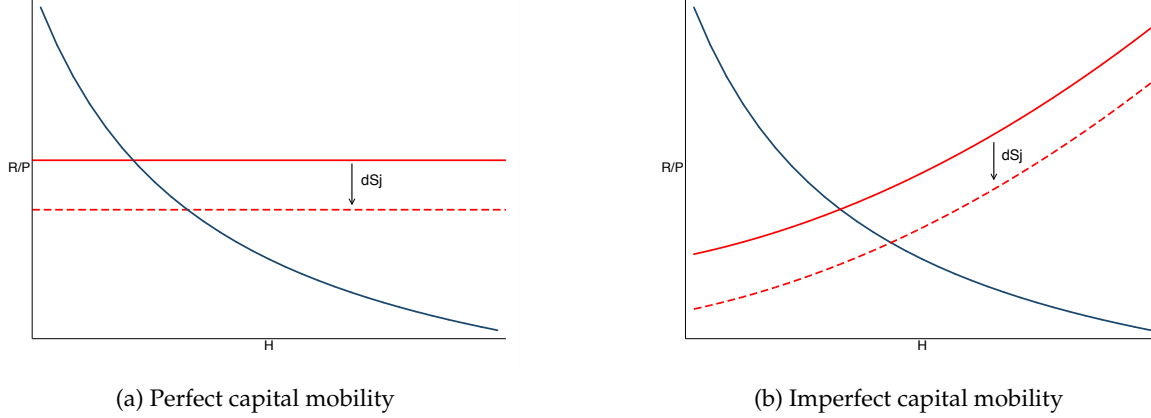
**Proposition 3.2.** (i) An increase in the subsidy  $dS_j > 0$  to the provision of rental services in location  $j$  raises housing production  $H_j$ , renter population  $L_j^W$ , and the price of both land ( $r_j^T$ ) and new dwellings ( $P_j^H$ ) in  $j$ ; it also increases net-of-tax returns to investment  $r_j^H = \frac{R_j S_j}{P_j^H}$ .

(ii) For  $\phi \frac{r_k^\phi}{\bar{m}_{Hi}^\phi + r_k^\phi} < \theta$ , the partial equilibrium effect of the subsidy  $\frac{\partial \log P_j^H H_{i \rightarrow j}}{\partial \log S_j} = s_{ij} \phi (1 - HS_i) + (1 - s_{ij}) \theta$  is stronger for investment flows  $H_{i \rightarrow j}$  stemming from origin locations  $i$  with a lower initial ownership share  $s_{ij}$ .

(iii) The response of overall local housing provision to subsidies  $\frac{\partial \log P_j^H H_j}{\partial \log S_j}$  is larger in treated locations with a lower share of self-ownership  $s_{jj}$ .

- *Effects of investment tax credits on investment, housing production, prices, and population* A location-specific subsidy  $S_j$  for the purchase of rental property in  $j$  shifts the allocation of savings towards housing in general, and location  $j$  dwellings in particular. The subsidy raises housing production  $H_j$  and sales  $P_j^H H_j$  in  $j$ , leading to higher land ( $r_j^T$ ) and new unit prices ( $P_j^H$ ). The induced increase in the housing stock triggers inwards renter mobility, inducing population growth  $dL_j^W > 0$  in treated locations.
- *Incomplete capitalization* The usual view of exogenously determined post-tax returns leaves no room for any incidence of place-based policies on savers. Under an imperfectly elastic capital supply, however, the subsidy is not fully capitalized in rents and prices. Post-tax returns  $\frac{R_j S_j}{P_j^H}$  in treated locations rise to induce new investors to switch to rental investment in  $j$  at the extensive margin. Therefore, the pre-tax return  $\frac{R_j}{P_j^H}$  drops by less than the full amount of the subsidy (either through higher acquisition prices or lower rents), and part of the incidence of place-based policies falls on infra-marginal landlords who would have invested in  $j$  even absent the subsidy and enjoy higher net returns as a consequence of it. Figure 7 evidences this phenomenon, showing how the outwards shift in the supply of capital fails to be fully capitalized in pre-tax returns under an upwards-sloping supply of capital.<sup>41</sup>

<sup>41</sup>The full capitalization under perfect capital mobility is reminiscent of the effect of investment tax credits on prices in House and Shapiro (2008). Due to inter-temporal substitution in investment, the ability to substitute capital goods acquisitions across *time* implies



**Figure 7: Imperfectly elastic capital supply and the effects of landlord subsidies** Figure 7 summarizes the role of imperfectly elastic capital supply for the response of housing investment and returns to landlord subsidies. Panel (a) presents a generic city in a financially integrated country: returns fully capitalize the value of the landlord incentive, leaving post-tax returns unaffected, either through a rise in prices or through lower rents. Panel (b) plots the case of an imperfectly elastic supply of capital: here, while the production of new housing also rises, part of the incidence of subsidies falls upon infra-marginal savers who receive higher post-tax returns.

- *Rise in long-distance investment* An implication of equation 3 is that the partial equilibrium elasticity of housing investment flows from  $i$  to  $j$  varies systematically with the portfolio share of  $j$  in  $i$ ,  $s_{ij}$ :

$$\frac{\partial \log P_j^H H_{i \rightarrow j}}{\partial \log S_j} = s_{ij} \phi (1 - HS_i) + (1 - s_{ij}) \theta$$

When the conditional share  $s_{ij} = \frac{(e^{-\delta d_{ij}} r_j^H)^\theta}{\sum_{j'} (e^{-\delta d_{ij'}} r_{j'}^H)^\theta}$  of  $j$  in  $i$ 's portfolio is high (i.e. when bilateral frictions  $d_{ij}$  are low), the partial equilibrium elasticity of investment to local returns is mostly governed by the reallocation from outside assets to housing, which depends on  $\phi$ . On the contrary, when  $s_{ij}$  is low (more distant locations), the partial equilibrium responsiveness of investment flows to returns mainly depends on  $\theta$ , the parameter that affects the ease of "within-housing", lower-nest reallocation. Intuitively, nearby investors are mostly infra-marginal: most choose to invest in  $j$  independently of the subsidy, and their response, if any, operates through the difficult reallocation of non-housing to housing assets. On the other hand, for long-distance investors, a shock to returns in  $j$  shifts their large non- $j$  housing investments towards the treated location.<sup>42</sup> With  $\theta > \phi(1 - \text{Housing share}_i)$ , a higher subsidy thus raises the proportion of long-distance investors.

- *Heterogeneous responses by access to outside capital* Aggregating flows from all locations, the response of

that firms' demand for long-lived assets is almost infinitely elastic to temporary changes in the tax treatment of investment, and the price of equipment goods should reflect the full value of tax incentives. Similarly, under the (violated) null assumption of perfect capital mobility across space, the ability of investors to perfectly arbitrage return differentials between subsidized and untreated locations would imply a demand for residential capital assets that is perfectly elastic to returns across places, and, in turn, full capitalization of the value of place-based incentives in prices.

<sup>42</sup>This phenomenon is similar to the findings of Berger, Herkenhoff, and Mongey (forthcoming), who study the response of employment to firm wages. When a firm has a large employment share in a given market, most of the response to higher wages comes from cross-market employment reallocation. Here, when housing in  $j$  represents a large portfolio share in  $i$ , the bilateral investment response to return shocks is drawn from cross-asset (rather than "within-housing") reallocation.

housing provision to a place-based subsidy is heterogeneous across space. Since the response is stronger for longer-distance investors as described above, the locations that benefit most from the subsidy are those with more dispersed initial ownership shares, where a larger baseline proportion of housing is owned by these more distant landlords.<sup>43</sup> Formally, this openness to outside capital is summarized by a model-implied measure of the dispersion of ownership:  $1 - \sum_i \frac{H_{i \rightarrow j}}{H_j} s_{ij}$ , the sum of  $i$  investors' housing portfolio shares in  $j$ , weighted by the baseline share of investors from  $i$  in  $j$ .<sup>44</sup>

## 4 Quasi-experimental evidence: setup, strategy, and data

To assess the causal impact of tax subsidies on the spatial allocation of residential investment, and show that they are consistent with the mechanism described in section 3, I turn to quasi-experimental evidence on landlords' decisions. I rely on exogenous place-specific changes in eligibility to a tax credit for buy-to-let investment, and estimate its consequences for the housing market, landlords, and local communities. This section describes the institutional setup I exploit for causal inference (4.1), defines my empirical strategy (4.2), and summarizes the data (4.3).<sup>45</sup>

### 4.1 Institutional background

**The Pinel housing tax credit** In January 2013, French Housing Minister Cecile Duflot introduced a conditional individual income tax rebate for taxpayers purchasing newly built real estate in eligible locations, if they offered affordable rents for nine years.<sup>46</sup> The *Duflot* law was revamped in September 2014 by Duflot's successor, Sylvia Pinel. Under the new *Pinel* tax credit, individual landlords purchasing a new dwelling receive a personal income tax reduction of up to 21% of their investment if they lease the unit for a minimum of six years below a regulatory rent ceiling, and tenant income does not exceed a threshold amount.<sup>47</sup> Rent and income ceilings both vary with the location of the municipality (see *infra*). Annual investment is limited to two dwelling units and a maximum of €300,000 (ca. USD 350,000) per year.<sup>48</sup> The reduction only applies up to a purchase price of €5,500/sq.m., creating a kink in the credit. Several changes made the *Pinel* scheme more favorable than the *Duflot* mechanism. A more flexible rental duration commitment (6, 9, or 12 years) was

<sup>43</sup>As in standard spatial frameworks, the heterogeneous response to location-based subsidies also depends on "physical" supply conditions arising from locally varying physical land abundance and intensity in the production function (denoted by the composite elasticity  $\frac{1+\sigma_j-\xi_j}{\xi_j}$ ).

<sup>44</sup>This measure is highly correlated with  $1 - s_{jj}$ , the complement of the own-ownership share in the housing stock. It is reminiscent of the heterogeneity in local employment elasticities in Monte, Redding, and Rossi-Hansberg (2018) that stems from variation in the "own-employment" share and makes cities more open to commuting more able to accommodate labor demand shocks.

<sup>45</sup>Appendix A includes additional information on the data sources used in the paper. Appendix B provides more detail on the institutional setup and French housing supply incentives.

<sup>46</sup>Relative to its predecessor, the *Scellier* scheme implemented in 2009, the *Duflot* law had lower levels of investment tax credits, and more stringent affordability requirements, with rent and income ceilings close to 20 percent lower (see appendix B.1).

<sup>47</sup>The tax reduction is computed as 2% of the purchase price each year after the start of the lease, for up to six or nine years (depending on the taxpayer's initial affordable rent duration commitment), with an option for a 1% additional annual reduction for years 10 to 12, if the taxpayer extends her affordability commitment - yielding a maximum reduction of 21%.

<sup>48</sup>The maximum subsidy is therefore  $0.21 \times 300,000 = 63,000$  euros over twelve years. The *Pinel* scheme is a non-refundable tax reduction. Because of the indivisibility of housing units documented in section 2, the 2 percent annual non-refundable reduction only fully benefits landlords with a substantial tax liability to offset: around 60% of beneficiaries are in the top decile of household income, and 90% in the top tercile, according to government reports (Inspection Generale des Finances, 2019).

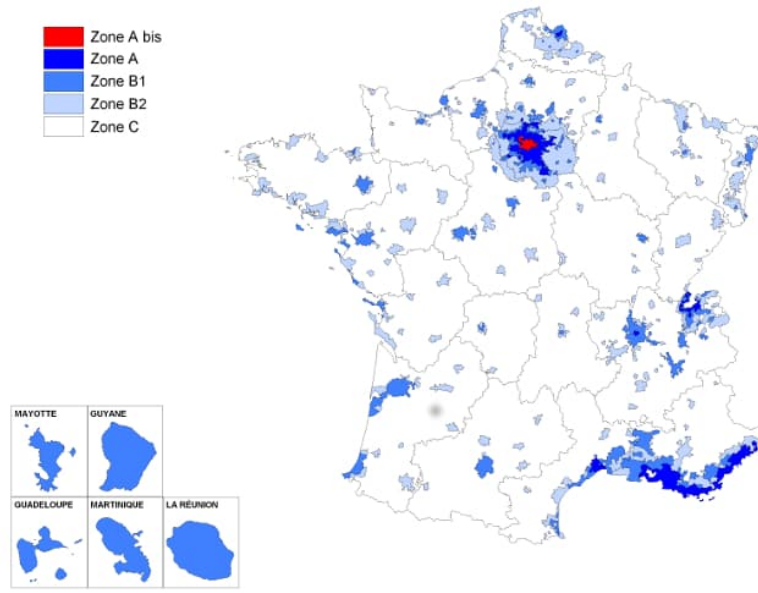


Figure 8: **Pinel zoning** Municipalities were mapped into one of five areas (Abis, A, B1, B2, C), determining both their eligibility to the *Pinel* scheme, and rent guidance and income restrictions. Areas are ranked in descending order of estimated property market “imbalances”. The classification was widely reshuffled with the implementation of the *Pinel* scheme in October 2014 Source: French Housing Secretariat.

available under the *Pinel* tax credit (only 9 years with the *Duflot* scheme); a prohibition to rent the unit to one’s relatives was removed; and a reshuffling of eligible areas (see *infra*) broadened access to the scheme. Table 3 provides a range for the net present value of the landlord incentive between 9 and 15 percent, accounting for varying interest rates and rent discount stringency estimates.

**Place-based targeting** Eligibility, maximum rents, and income ceilings under the *Pinel* scheme follow strict place-based regulations (known as the *zonage Pinel*) by municipality. Nationwide rules classify France’s 35,000 municipalities into five classes, ranked by descending order of government-assessed housing market “tightness”. These classes, illustrated in figure 8, range from *A bis* (Paris and its closest neighbors), to *A* (the broader Paris region, Marseille, Lyon, and some touristic towns on the French Riviera), *B1* (several other large cities, including Toulouse, Bordeaux or Strasbourg), *B2* (most medium-sized cities and their surroundings, like Dunkirk, Saint-Etienne, or Le Mans), and *C* (all 30,000 other towns). Under the scheme, each “zone” is subjected to different rent ceilings and maximum renter income. Rents per square meter are capped at  $R = M \times C$ , the product of a coefficient  $M$  (decreasing with dwelling size and capped at 1.2:  $M = \min[\frac{19}{\text{Size}} + 0.7, 1.2]$ ), and a zone-specific ceiling amount  $C$ .<sup>49</sup> To accompany the roll-out of the *Pinel* scheme, a September 2014 government ruling modified the allocation of municipalities to the five classes, leading almost exclusively to zone upgrades and thus to less stringent rent and income ceilings for municipalities representing about 10 million inhabitants, close to a sixth of the total French population.<sup>50</sup> As shown in appendix B.2, there is a wide heterogeneity in

<sup>49</sup> Appendix B.1 provides additional details on the computation of the cap for local rents and incomes.

<sup>50</sup> 162 B1 municipalities (with close to 4 million residents, including the major cities of Lyon, Lille or Marseille) and 25 B2 municipalities (approx. 0.2 million res.) were upgraded to A. 8 municipalities (approx. 0.2 million res.) were upgraded from A to A bis. 358 B2

incomes and rents among municipalities *within* a zone, and a significant overlap in housing market conditions and demographics *across* them. Nonetheless, applicable rent and income ceilings are uniform within a *Pinel* zone, and differ across them, yielding some variation in effective local stringency.<sup>51</sup>

**Policy timing and reforms** The *Duflot* tax rebate ran from January 2013 to August 2014, followed by the *Pinel* law after September 2014. Both policies were available, *de jure*, for the purchase of any new unit located in the A, A bis, and B1 areas. Units in C towns were not eligible to either scheme. Eligibility to the *Duflot* law in the approximately 3,900 B2 towns was phased out after June 2013. However, B2 towns could request an agreement from regional authorities in order to participate in the scheme beyond June 2013.<sup>52</sup> Authorizations were granted to around 900 B2 municipalities,<sup>53</sup> whose property markets were assessed to be closer to (always eligible) B1 municipalities. A purchase in an “agreed” B2 municipality (henceforth a “B2+” town) was thus eligible to the *Duflot* incentive from the date of the agreement until August 2014, and to the *Pinel* credit afterwards, while purchases in other B2 cities (henceforth a “B2-” town) were never eligible to the *Pinel* subsidy. The *Pinel* scheme was initially due to expire on December 31<sup>st</sup>, 2016, but was extended in late 2016 until December 31<sup>st</sup>, 2017. In December 2017, the French parliament extended the scheme again until December 2021, but restricted it to areas A, A bis, and B1 to reduce its cost, thus excluding B2+ municipalities.<sup>54</sup> To smooth the transition for construction projects already under way in B2+ cities losing the benefit of the policy, sales in the B2+ area remained eligible to the tax break until December, 31<sup>st</sup> 2018 (later extended for technical reasons to March 15<sup>th</sup>, 2019), as long as the building permit had been requested before December 31<sup>st</sup>, 2017.<sup>55</sup>

## 4.2 Empirical strategy

### 4.2.1 Two natural experiments: entry and exit

*Pinel* scheme subsidies varied over time with a location’s eligibility to the policy. Nationally mandated changes in coverage by the policy generate sharp variation in the tax treatment of residential investment, plausibly unrelated to local demand conditions or government priorities – a place-based setting that differs from a large literature focused on the role of tax or subsidy competition among communities deliberately offering different

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municipalities (hosting close to 3.4 million residents) and 167 C municipalities (approx. 0.7 million res.) were upgraded to B1. 462 municipalities (approx. 2.4 million res.) were upgraded to B2. Conversely, only 38 municipalities (40,000 residents) were downgraded from A to B1, and 619 small municipalities (0.6 million res.) were downgraded from B1 to B2. No town was downgraded to C.

<sup>51</sup>Table 2 and B.2 provide further summary information on the distribution of population, demographic characteristics, rents, and incomes across municipalities by zone and agreement status.

<sup>52</sup>The competent authority was the prefect, the State’s representative in each of France’s 13 regions. Prefects would grant authorizations based on their qualitative appreciation of “tightness” in the local housing market.

<sup>53</sup>Close to sixty percent (530) of these agreements were put in place in 2013; more than thirty-five percent were granted in 2014 and 2015, with the remaining agreements given after 2016. 20 (out of a bit more than 30,000) municipalities in the C area also obtained a special agreement. A handful of municipalities gained eligibility when a non-eligible town merged with an eligible one.

<sup>54</sup>The costs of tax breaks for landlords or corporate developers makes them a regular target of budget-balancing proposals, both in Europe and in the United States: the Congressional Budget Office regularly mentions “Repeal the Low-Income Housing Tax Credit” ([link](#)) as an effective revenue-enhancing avenue in its annual *Options for reducing the deficit*.

<sup>55</sup>At the same time, the policy was expanded, starting January 2018, to cover the so-called CRSD (*Contrats de Redynamisation des Sites de Defense*) towns hosting former defence bases and receiving transitional financial support from the government when military installations closed. Therefore, a few B2+ municipalities with an agreement *remained* eligible to the policy, a feature I take into account when estimating the effect of the removal of the policy. Other B2- or C towns which had not hitherto been eligible *gained* access to it. I assess the effect of the subsidy in formerly ineligible (B2-/C) CRSD towns in appendix B.3.

location incentives (Slattery and Zidar, 2020).<sup>56</sup> However, eligible A, A bis, B1 and B2+ municipalities were characterized by distinct housing prices and overall market trends, than untreated B2- and C towns, since the classification was constructed to correspond to relative property market tension (as shown by the starkly different characteristics of each group summarized in appendix B.2 and table 2). Comparing the evolution of outcomes for all eligible and all ineligible towns would require strong parallel trends assumptions unlikely to be met, and would not disentangle the effect of subsidies from substantial composition differences in local dynamics. Fortunately, the government-designed partitioning of French municipalities into areas with comparable housing market trends builds natural control groups, for which differential dynamics in outcomes over time can credibly be attributed only to diverging subsidy eligibility. Specifically, I exploit two natural experiments linked to the timing and place-targeting of the policy:

1. **Loss of eligibility (*exit-DiD*)** To assess the short-run effect of tax credits on buy-to-let investment and eligible home sales and prices, I use a difference-in-differences strategy exploiting the *end of eligibility* for B2+ towns in December 2017, comparing formerly treated B2+ towns to B1 municipalities that remained eligible after the policy was scaled down. The exogenous differential loss of eligibility allows for a direct estimate of the response of residential investment by individual landlords. I label this difference-in-differences around the timing of eligibility loss the “*exit-DiD*” approach.
2. **Entry into eligibility (*entry-DiD*)** To explore medium-term consequences of tax subsidies for the broader local real estate market, housing conditions, and local communities, I estimate a difference-in-differences model around the timing of *entry* at the start of the *Pinel* scheme, leveraging the roll-out of agreements to B2+ municipalities, and comparing them only to a large permanent control group of never-treated “B2-” municipalities. I label this study of entry into policy eligibility the “*entry-DiD*” approach.

#### 4.2.2 Preferred specifications

**Difference-in-differences around loss of eligibility** I estimate the short-term effect of removing access to the *Pinel* tax credit, using the *exit-DiD* approach. In this specification, I take advantage of the exogenous loss of eligibility in B2+ towns decided by the national government to directly estimate two outcomes of interest: the residential investment response, and the price effects of losing the subsidy. Since municipalities in the B1 area remain treated after the 2017 reform, while B2+ towns were no longer eligible, the *exit-DiD* specification restricts the sample to B2+ and B1 municipalities only. For an outcome  $y_{ct}$  in municipality  $c$  and time period  $t$ , I estimate a full leads-and-lags regression in the window  $[\underline{T}, \bar{T}]$  leading to and following the end of eligibility:

$$y_{ct} = \alpha_c + \gamma_t + \sum_{\substack{k=\underline{T} \\ k \neq -1}}^{\bar{T}} \beta_k \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t=k} + \Gamma X_{ct} + \epsilon_{ct} \quad (6)$$

<sup>56</sup>The open-ended nature of the French buy-to-let tax credits, conditional on zone eligibility, is reminiscent of the Opportunity Zones implemented in the United States by the TCJA in 2017. This departs from schemes predicated on fixed local allotments of tax credits, such as the LIHTC. Incentives embedded in the LIHTC have been hard to estimate empirically, since a fixed amount of credits allocated to states are awarded as grants to projects at the discretion of local authorities. Quasi-experimental approaches to estimate the investment response have relied on a cross-sectional discontinuity in the LIHTC formula (Baum-Snow and Marion, 2009), which depends on the proportion of low-income households in a Census tract, and defines Qualified Census Tracts and Difficult Development Areas.



where  $\gamma_t$  indexes a period fixed effect accounting for common time trends,  $\alpha_c$  is a municipality fixed-effect that averages out municipality-level, time-invariant determinants,  $X_{ct}$  is a vector of time-varying controls at the municipality level (including initial municipality characteristics interacted with time fixed-effects), and the coefficients  $\beta_k$  measures the time-varying impact of eligibility removal on outcomes. The end of eligibility is the date after which units in B2+ locations were no longer eligible – corresponding to early 2019 for sales, and to December 2017 for permit requests. The key identification assumption is that, conditional on other characteristics, B2+ municipalities would have followed parallel trends to *always-eligible* B1 cities after that time, were it not for losing *Pinel* eligibility, since B2+ cities were deemed by regional prefects to more closely fit into the B1 classification. The leads-lag specification allows me to graphically visualize any differential pre-trends at high frequency *before* the removal of eligibility.

**Difference-in-differences around entry** After providing evidence on the immediate impact of losing eligibility to the policy for B2+ cities, I turn to its longer-run effects by exploiting the differential *entry* of B2+ cities – relative to B2- – after October 2014. I first confirm the quantitatively similar effect on sales at the intensive and extensive margin, before showing how the local outcomes for long-distance investment, the housing stock, and renter mobility track entry into eligibility among B2+ cities receiving an agreement, relative to a permanent control group of never-treated B2- locations. In this *entry-DiD* approach, I restrict the sample to all B2 municipalities in the period lasting until the end of eligibility, and use the following specification:

$$y_{ct} = \alpha_c + \gamma_t + \beta_R \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t \geq \text{Entry start date}} + \Gamma X_{ct} + \epsilon_{ct} \quad (7)$$

in order to estimate the effect of entering eligibility to the subsidy on posterior outcomes for municipality  $c$  at time  $t$ . I also use a fully dynamic specification of the form:

$$y_{ct} = \alpha_c + \gamma_t + \sum_{\substack{k=\bar{K} \\ k \neq -1}}^{k=\bar{K}} \beta_k \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t=t_0(c)+k} + \beta_{\underline{K}} \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t < t_0(c)+\underline{K}} + \beta_{\bar{K}} \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t > t_0(c)+\bar{K}} + \Gamma X_{ct} + \epsilon_{ct} \quad (8)$$

where  $\mathbb{1}_{t=t_0(c)+k}$  equals 1 for treated municipalities if calendar time period  $t$  is the  $k^{\text{th}}$  period before or after  $t_0(c)$ , the eligibility start date for town  $c$  (the start of the *Pinel* scheme), and 0 otherwise.<sup>57</sup> This specification allows me to visualize any differential trends prior to gaining eligibility, giving credence to the untestable assumption of parallel post-entry trends. The underlying identification assumption is that, in the continued *absence* of an agreement, B2+ cities would have followed comparable trends to *never-eligible* B2- towns. The government’s decision to classify B2 towns in a common zone suggests they initially had comparable observable dynamics. Time-varying controls, sub-sample analyses, and robustness checks support the credibility of the identification assumption.

<sup>57</sup>The continued ineligibility of non-agreed “B2-” municipalities creates a natural permanent and large control group of never-treated cities, a feature particularly relevant to the identification of the dynamic path of treatment effects (Borusyak, Jaravel, and Spiess, 2021).

### 4.2.3 Potential threats to identification

**Selection into treatment** Specific to the *entry* approach, one may be concerned about selection of B2+ towns relative to B2- cities, if locations expecting stronger investor demand were more likely to request and/or obtain an agreement. To emphasize the causal role of tax subsidies in triggering additional investment at the local level, I evidence the absence of differential pre-trends between eligible B2+ and untreated B2- areas, prior to eligibility changes. I also provide within-municipality-month placebo tests, by comparing the dynamic impact of eligibility on sales of new homes, to its placebo and null effect on sales of (ineligible) existing units. Sharp trend breaks at the beginning and end of eligibility, and high-frequency time bunching in anticipation of policy changes, also do not fit alternative explanations based on lower-frequency differential trends, and suggest a causal effect of the policy. Finally, the *exit-DiD* approach, in which the loss of eligibility is exogenous and shown to trigger immediate trend breaks in local outcomes, is an additional check allowing me to rule out the selection explanation for differential outcomes in B2+ towns during the earlier period of policy eligibility.

**Other place-based government policies** The validity of causal estimates obtained from the specifications above may be threatened by other time-varying determinants of municipal level-outcomes (included in  $\epsilon_{ct}$ ) evolving jointly with the spatial allocation and the time path of eligibility, chief among which is the presence of other place-based government policies. The only case of such a potential confounder is France's interest-free loan policy for first-time homebuyers of owner-occupied units (*Pret a Taux Zero* or PTZ). The generosity of the interest-free loan policy varied with zoning, and was amended specifically for B2 and C areas after 2017.<sup>58</sup> Two aspects of my empirical strategy help assuage this concern. First, using the differential treatment of B2+ and B2- municipalities upon *entry* provides compelling evidence that any effects are driven by the *Pinel* policy, since PTZ eligibility and amount do not vary between these two subsets of B2 towns. Second, I show that the impact of the *Pinel* policy is concentrated on buy-to-let units, which are ineligible to the PTZ, and high-frequency effects are only consistent with *Pinel* variation.

## 4.3 Summary of the data

I summarize the key relevant data sources used to examine the impact of the *Pinel* scheme on build-to-let sales, prices, the housing stock, local communities, landlords, and renters.<sup>59</sup>

**Housing markets** I avail myself of exhaustive data recording all housing and land transactions in France from January 2010 to December 2019, recorded by the French Treasury for tax purposes in the *DV3F* database, and comprised of 800,000-1,000,000 transactions each year over the period. I also use an exhaustive survey of developer-led projects, the *Enquete sur la Commercialisation des Logements Neufs* (ECLN), recording detailed

<sup>58</sup>After January 1<sup>st</sup>, 2018, the PTZ, a policy providing households with interest-free mortgages when they buy their first owner-occupied home, was restricted to 20 percent of the cost of first-time home purchases in the B2 and C area, down from 40 percent.

<sup>59</sup>Beyond the main restricted-access sources mentioned here, some surrogate sources exploited in the stylized facts characterized by section 2, and already mentioned then, are not described in detail here, but more information is provided on the construction of all sources in appendix A.

information on all new housing sales by developers at the dwelling level.

**Housing stock** I use the *Sit@del2* database from the French Housing authority, comprising the universe of housing permits requested from 2009 to 2019. To study changes in the make-up of local dwellings, I rely on register data - the *Fichiers Fonciers*, an annual cadaster of all occupied and vacant housing units in France. I also use the *Répertoire des logements locatifs des bailleurs sociaux*, a repository of all social housing units in France, to study the extent of crowding-in or crowding-out of social housing by intermediate-rent housing construction incentivized through the *Pinel* scheme. The *Observatoire de l'artificialisation des sols* database, which combines satellite information with administrative records at the parcel level, allows me to compute the amount of land converted each year from "natural" use to residential or industrial use, in each municipality.

**Investor behavior and ownership structure** I make use of the FILOCOM/FIDELI database, a repository of close to thirty million housing units (as of 2019) serving as primary residences in France, to measure the evolution of the ownership network. I also use the universe of French individual tax returns in 2016, an exhaustive sample of all tax units each year (collectively known as *Fichiers POTE - panelisables*), to measure landlord status and compute rents and purchase prices for subsidized units.

**Local communities** Data on municipal population, as well as municipality-to-municipality annual migration flows, assembled by the French statistical institute INSEE through the Census, are used to estimate population mobility. To look at specific mobility effects among lower-income households who rent, I rely on data from the *Caisse d'Allocations Familiales*, France's main outlet for social benefit payments, recording the annual number of beneficiaries of housing allowances, as well as the number of low-income households in each municipality. When assessing whether the policy shapes local economic activity, I exploit establishment-level information from the *Repertoire des Entreprises et des Etablissements* on the number of employees and exact location of each establishment in France, aggregated at the municipality-year level.

## 5 Housing policy with home-biased landlords: results

This section analyzes the direct and indirect consequences of the place-based *Pinel* rental investment subsidy, and shows that its effects on landlords, local real estate markets, and residents, are consistent with the conceptual mechanism of imperfect capital mobility across space advanced earlier.

### 5.1 The impact of place-based subsidies on housing investment

**The impact of losing eligibility to the subsidy** I first examine the short-run implications of losing eligibility to the *Pinel* subsidy on purchases of eligible new dwellings from developers by individual investors, exploiting the quasi-experiment induced by its discontinuation in B2+ cities after 2018. Validating the causal policy channel and the first part of proposition 3.2, figure 9 displays the key result. It plots the quarterly coefficients  $\beta_k$

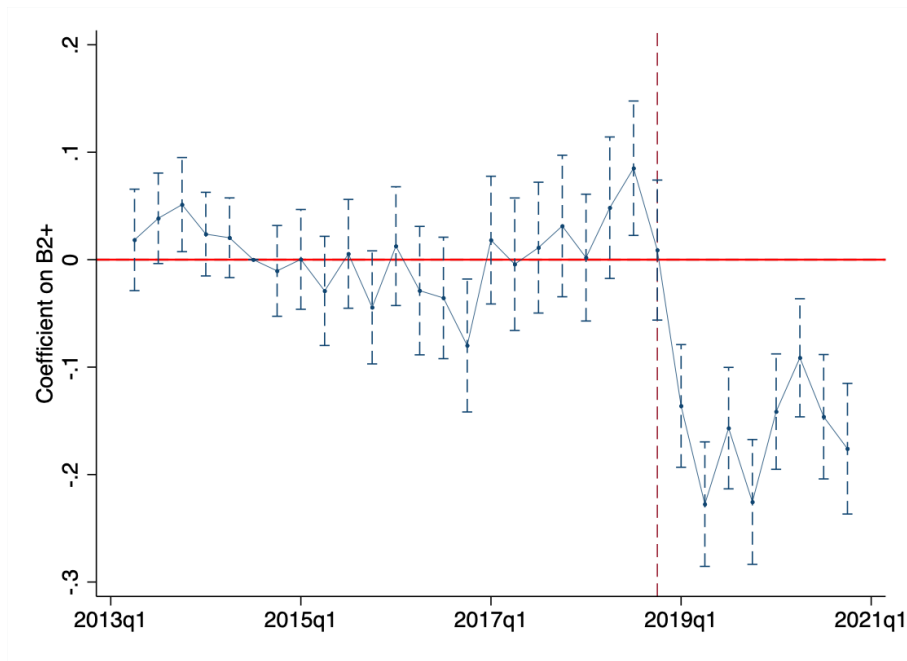


Figure 9: **The impact of exiting the subsidy on new home sales** Figure 9 documents the estimates of coefficient  $\beta_k$  in equation 6 for the impact of losing eligibility to the subsidy (in B2+ cities) on (log) sales of new buy-to-let units by developers, relative to B1 (always-treated) cities. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are clustered at the municipality level, and 95% confidence intervals are shown in brackets. The dashed vertical line marks the end of eligibility of B2+ towns to the *Pinel* incentive relative to B1 towns.

on a time-varying dummy for being part of the B2+ group, in the full leads-and-lags specification of equation (6). The dependent variable is the (log) number of buy-to-let units sold in a municipality each quarter, and the sample is restricted to B1 (always-treated) and B2+ (losing eligibility) towns. Quarterly sales follow precisely parallel dynamics throughout the duration of common eligibility to the policy, with no evidence of differential pre-trends. When the policy is discontinued, after an initial spike in the last quarter in the anticipation of its end, the number of sales drops by about twenty percent in B2+ municipalities relative to B1 towns, an economically and statistically significant impact with no apparent recovery up to eight quarters later.<sup>60</sup>

I assess alternative specifications and the robustness of the result in figure ii and table 5. First, to show that the result is not driven by functional form choices or regression adjustments, panel (a) of figure ii plots the (normalized) quarterly number of new buy-to-let units sold by developers in two specific subsets of comparable municipalities, all classified as B2 until the September 2014 re-zoning: those later upgraded to B1 by the rezoning (“B2 to B1”, eligible to the *Pinel* scheme throughout the 2014-2020 period) and those that remained in the B2 group with an agreement (“B2 to B2+” towns, eligible only until December 2018). Sales of new units rose in tandem in both groups after the 2014 introduction of the *Pinel* investment tax credit, reflecting their common eligibility. However, while they remained at an elevated level in the “B2 to B1” group of “always-treated” towns, they dropped sharply and immediately once eligibility was discontinued after 2018 in “B2 to B2+” towns.

<sup>60</sup>All specifications cluster standard errors at the level of individual municipalities, to reflect potential serial correlation in error terms and the level of treatment assignment.

Panel (b) of figure ii plots the quarterly coefficients  $\beta_k$ , where the dependent variable is the total number of units sold in a municipality each quarter and the sample is again all B2+ and B1 towns. Sales of overall new units drop after the end of the policy, although by a more limited magnitude than the specific effect on buy-to-let eligible dwellings, as predictable if some units are sold to owner-occupiers or institutional investors. Panel (c) in figure ii shows the effect on the (log) value of sales of eligible dwellings. Finally, panel (d) examines the extensive margin response of housing investment at the town level. It plots quarterly coefficients on the B2+ dummy in a specification where the outcome is the presence of *any* new home sale in the municipality in a quarter. There is an 8-10 percent drop in the probability of a sale occurring in formerly treated B2+ towns after they lose the tax-favored treatment of new rental dwellings.

All specifications (summarized in table 5) evidence closely parallel trends (with pre-exit coefficients statistically indistinguishable from 0) for the B1 and B2+ groups throughout the period of joint eligibility, from October 2014 to December 2018; and a sharp trend break after B2+ towns lost eligibility to the *Pinel* subsidy. Overall, the results highlight the positive but quantitatively limited user cost elasticity of residential investment by individual landlords. The summary computation in table 3 estimates the net present value of the tax credit, once accounting for rent discounts and depending on assumptions on the discount rate and baseline price-to-rent ratios, ranges from 9 to 15 percent of the purchase price of a unit. Therefore, the estimated reduced-form effect is consistent with a user cost elasticity in the range of 1.3 to 2.1, markedly lower than firm-level estimates in Zwick and Mahon (2017).<sup>61</sup>

**The impact of gaining eligibility** I next turn to the *entry* approach, comparing B2+ towns (that benefited from the *Pinel* scheme from 2014 to 2018) to B2- towns, which, although part of the same class, were never granted eligibility to the incentive. Figure 10 plots the estimated time-varying causal effect of being part of the B2+ group on eligible new home sales to investors (equation 8), in the restricted sample of (eligible) B2+ and (untreated) B2- towns over the 2013-2020 period. After the implementation of the *Pinel* scheme, sales increase in eligible towns, and display a clear divergence from B2- control towns. The intensive margin impact of 20 to 25 percent at the peak is in line with the estimated effects of *exiting* the policy plotted in figure 9.

I present alternative specifications in figure iii: the total value of sales in the B2+ and B2-; and the coefficients from estimating equation 8 for (log) sales of all units, and the intensive and extensive margins responses of any new sale to rental investors – respectively in panels (a), (b), (c), and (d). Panel (a) provides aggregate evidence that the effect of the policy upon entry was noticeable: sales dynamics in B2+ towns exhibit a sharp increase after the start of the scheme, while sales in B2- cities remained mostly flat throughout the duration of the policy. In each of the other three panels, the time-varying quarterly coefficient  $\beta_k$  on being part of the B2+ group is statistically indistinguishable from zero both before and after the differential policy eligibility period in the sample of all B2 towns, but B2+ towns exhibit a temporary departure from trend corresponding exactly to the period of availability of the investment incentive. Estimated magnitudes at both the intensive and extensive margins are summarized in table 6. They are quantitatively consistent with the estimated effect

<sup>61</sup> Appendix B.1 provides a discussion of the assumptions and sensitivity analyses for the implied net present value of the *Pinel* incentive, for a range of parametrizations of the discount rate, and bite of affordable rent markdowns.



Figure 10: **The impact of gaining eligibility to the subsidy on new home sales** Figure 10 documents the estimates of coefficient  $\beta_k$  in equation 8 for the impact of gaining eligibility to the subsidy (in B2+ cities) on (log) sales of new buy-to-let units by developers, relative to B2- (never-treated) cities. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are clustered at the municipality level, and 95% confidence intervals are shown in brackets. The two dashed vertical lines mark (respectively) the beginning and end of differential eligibility of B2+ towns to the *Pinel* incentive relative to B2- towns.

of *losing* eligibility examined *supra*.

Therefore, *entry* into eligibility to the *Pinel* policy among B2+ towns obtaining an agreement is associated with similar but opposite-sign effects to their later exit out of the policy. Beyond providing an alternative experiment to test the robustness of the findings on new dwellings sales, the similar magnitudes give reassurance that B2- cities, a never-treated permanent control group, also provide an accurate counter-factual for the trajectory that other outcomes would have followed in eligible B2+ towns, absent the subsidy.

**A placebo test: existing home sales** The impact of eligibility to the policy may nonetheless be confounded by differential local trends if B2+ cities exhibit systematically different housing market activity from B2- towns during the period. Using administrative records of transaction deeds from the *DV3F* database, I can compare the relative evolution of new home sales (eligible to the *Pinel* tax credit) and *existing* home sales, which are never subject to the landlord incentive.<sup>62</sup> Contrary to sales of new homes, which increase substantially in B2+ locations (figure iv, panel (a)), transactions for existing homes (panel (b)) are unaffected and follow exactly parallel trends in B2+ and B2- towns throughout the duration of the policy. This within time-city placebo test shows that the increase in developer sales in treated locations during the eligibility period is not the product of differential housing turnover, which would lead these towns to experience higher sales activity overall. On the contrary, the null effect on sales of ineligible, existing dwellings confirms the causal interpretation of the

<sup>62</sup>New homes are identified in the notary deeds data by their value-added tax treatment, since existing homes are not subject to the VAT.

differential increase in new home sales, and supports the implicit parallel trend assumptions.

**Real effects on new construction** The previous results demonstrate that place-based investment tax credits substantially increase new home sales in eligible areas during the policy period. However, this behavioral response may reflect the crowding-out of un-subsidized rental housing production or owner-occupied units; tear-downs and renovations of existing dwellings; or simply faster outflows of developer inventory, rather than an effective new supply of local housing.<sup>63</sup> To show the limited extent of crowding-out effects, and evidence that the *Pinel* landlord incentive indeed increased the provision of housing space, I next show that the policy, rather than merely leading to faster outflows of existing developer inventory, triggered entirely new housing starts for multi-family buildings. Exploiting the universe of individual building license requests from 2010 to 2020, I examine building license requests, as well as authorized new units and housing starts in targeted B2+ areas relative to excluded B2- towns, during the policy period. Figure v summarizes the results.

Its top panels plot the quarterly number of authorized multi-family units in B2+ (once-treated) and B2- (control) or B1 (always-treated) municipalities, normalizing counts by the pre-period average quarterly number of authorizations. Licenses granted in B2+ and B2- municipalities closely track each other before the start of the policy, even at high-frequency seasonal variation. After the policy starts, however, the number of licenses granted takes off in eligible cities relative to the trend in ineligible cities, with a noticeable divergence throughout the year 2017, up until December 2017. Sharp spikes occur in the months surrounding the end of construction eligibility in December 2017, a natural consequence of time-bunching in licenses granted to corporate developers evidences earlier (see appendix D.1 for additional evidence on such time-bunching). After eligibility ends for building license requests in B2+ towns, the trend in these formerly eligible locations rapidly falls back to the corresponding path for ineligible cities. The two series come back in lockstep as the differential subsidy to eligible areas ends, and exhibit similar trajectories in 2019-2020. This differential variation in authorized units, closely mirroring the timing of policy eligibility to the subsidy, implies that the *Pinel* subsidy was indeed responsible for the rise in authorized constructions in eligible municipalities.

Panel (c) confirms this pattern using the cumulative monthly number of *housing starts* in both types of B2 municipalities. Housing starts follow a similar qualitative pattern to housing units authorized. After following precisely parallel trends prior to the start of the *Pinel* scheme in September 2014, housing starts diverge shortly after the subsidy is put in place in eligible B2 cities, relative to their ineligible counterparts. Panel (d) estimates the full leads-and-lags specification for the evolution of licenses granted in B2+ cities relative to B2- locations at annual frequency, finding a close to ten percent increase in licenses at the 2017 peak.

**Urban sprawl and land conversion** New home construction may occur either through the demolition and replacement of the existing stock of housing; through increased density of new floor area built per square footage of land in already urbanized spaces; or through expansion of the urban area into formerly un-built

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<sup>63</sup>For example, Berger, Turner, and Zwick (2020) find that the First Time Homeowner Tax Credit in the United States failed to causally trigger any additional new construction, and only led to a reshuffling of the existing housing stock. Several studies of the LIHTC also find almost full crowding-out of unsubsidized construction by subsidized units.

plots of land, leading to urban sprawl. To evidence the impact of the policy on urban sprawl and the share of local land that is built-up, I avail myself of data combined by the French *Observatoire de l'artificialisation des sols*, reporting detailed measures of land use at the local level, notably the annual flow of land area converted from natural uses (forest or agricultural land) to new habitat. Figure vi shows the annual flow of land transformed from natural uses to residential uses ("artificialization"), across municipalities, depending on their eligibility status. It evidences differential dynamics in Pinel-eligible and ineligible areas, with a pronounced divergence occurring in 2016, when conversion of land for residential use picked up in B1 and eligible B2 municipalities, but remained low in ineligible B2 and C towns. The differential spike in 2016 is consistent with the timing of the rise in construction observed in eligible areas, relative to ineligible towns. The policy therefore led not only to increased new construction of housing in targeted locations, but also to substantial urban sprawl and conversion of land from natural to residential uses at the boundary of urban areas, a potential driver of increased commuting times and adverse environmental consequences.

**Aggregate effects of landlord subsidies on sales** In line with the model-consistent mechanism of non-housing to housing investment reallocation, I provide suggestive evidence that aggregate rental investment by individual owners responds to time-varying tax incentives, exploiting longer-run aggregate time series data and all return shocks induced by a more or less favorable tax treatment. Even before the *Pinel* scheme, French governments have provided tax subsidies to the individual purchase of newly built rental housing, starting in the 1990s – generally conditional on setting rents at an intermediate level between market-rate and public housing. Table 4 summarizes the successive schemes in operation from 1996 to 2021, noting whether the implementation of a new scheme resulted in increased or decreased support relative to earlier tax subsidies.<sup>64</sup>

Figure 11 shows new residential units purchased each quarter closely track the implementation of more generous subsidy schemes, suggesting a substantial aggregate responsiveness of rental investment to tax incentives.<sup>65</sup> Periods in which incentives were altered to be more favorable to investors all coincide with a sharp rise in the number of units sold by developers. Conversely, intervals with lower subsidies are accompanied by reduced sales.<sup>66</sup> The generosity of these successive tax credits reflects changing government priorities and endogenously responds to the state of the macro-economy, therefore requiring the difference-in-differences approach exploited in the core quasi-experiment of the paper to make causal claims. Nonetheless, while some of the variation, such as the 2009 trough, reflects business cycles regularities, high-frequency changes precisely track the evolution of tax credits, are oftentimes at odds with the macroeconomic cycle,<sup>67</sup> and strongly suggest the presence of aggregate effects of landlord subsidies on investment. The large observed rise in sales after the

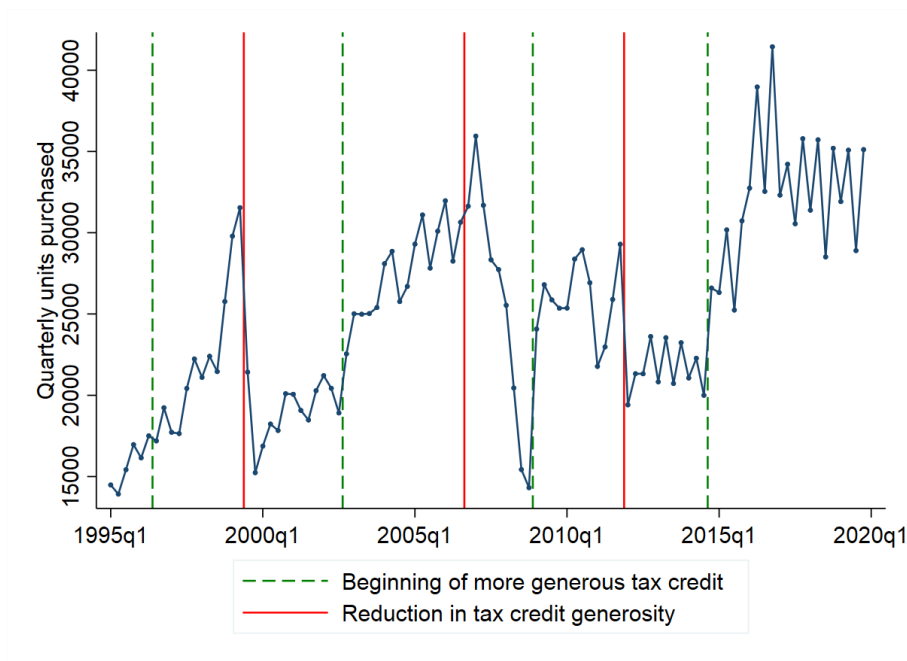
<sup>64</sup>Appendix B.1 provides additional descriptive information on the bonus depreciation and investment tax credits subsidizing new dwellings purchases by individual investors since 1996. Some resulted in a higher net subsidy to rental investment: the 1996-1998 initial *Perissol* accelerated depreciation scheme, the 2003-2005 mostly unconditional *Robien* tax credit, the 2009-2010 stimulus-motivated *Scellier* Act, and, finally, the 2014 flexible-duration *Pinel* scheme). Others reduced the net benefit to landlords: the 1999-2002 *Besson* law slowing down accelerated depreciation provisions, the 2006-2009 "re-focused" *Robien* scheme, the 2011-2012 budget-related cuts to the level of the *Scellier* tax credit, and the 2013 *Duflot* law with its more stringent affordability conditions.

<sup>65</sup>Figure E.6 plots the corresponding time series for the inflation-adjusted euro *value* of new residential investment.

<sup>66</sup>For the shorter, post-2010 period in which direct information on the final use of the asset is available, figure E.7 additionally evidences that the aggregate share of buy-to-lets in new units sales mirrors the time variation in subsidy generosity. It sharply drops at the end of the more generous version of the *Scellier* scheme in 2011, and rises after the implementation of the *Pinel* tax credit.

<sup>67</sup>e.g. during the dot-com boom and bust, or after the 2014 implementation of the *Pinel* scheme





**Figure 11: New units sold by developers, quarterly** The figure plots the total quarterly number of new housing units purchased from developers in France from 1995 to 2019. Dashed vertical green lines indicate the beginning of a more generous investment tax credit or bonus depreciation scheme for new housing; continuous vertical red lines indicate a switch to a less generous tax regime, either through reduced tax credit levels, stricter affordability requirements, or a deceleration of bonus depreciation schedules. The corresponding successive tax regimes are described in detail in table 4 and appendix B.1. The *Pinel* scheme starts with the last green line, in 2014Q4. The total number of units sold is computed from microdata in the *ECLN* database, an exhaustive survey of developer-led projects of five or more units.

implementation of the *Pinel* scheme in October 2014, from c. 20,000 quarterly sales to c. 35,000, is consistent with the substantial place-specific causal effects estimated earlier not cancelling in the aggregate.

Moreover, official deeds data allow me to document anticipation effects and inter-temporal tax arbitrage in aggregate sales. Figure vii documents the high-frequency pattern of actual signing dates for all new homes sales deeds recorded by notaries. Substantial time-bunching occurs in the weeks immediately preceding an anticipated shift to less favorable tax incentives. For example, while sales made before December 31<sup>st</sup> 2011 were eligible to the 2011 Scellier subsidy of 22 percent (for units respecting environmental guidelines – 13 percent otherwise), sales posterior to that date were only granted a lesser tax credit of 13 percent (6 percent for non-energy efficient units). Consequently, a marked bunching of sales occurs in the last days of 2011, immediately before the switch to the less generous subsidy.<sup>68</sup> Similarly, end-of-fiscal year spikes in recorded deeds are visible after the implementation of the *Pinel* subsidy scheme in 2014. This abnormal end-of-year concentration of transactions is likely triggered by increased certainty about a household’s precise income tax liability (and therefore the ability to collect the full non-refundable *Pinel* tax credit), as in Xu and Zwick (*forthcoming*)’s real option model of fourth-quarter corporate investment spikes.

<sup>68</sup>A tolerance period until March 2012 (for sales agreed with developers before December 2011) is also accompanied by a visible spike in deeds recorded immediately before March 31<sup>st</sup>, 2012, and a missing mass of sales in the following weeks. This anticipation result confirms the findings of Singh (2019), who documents that developers bunch residential investment decisions immediately before eligibility deadlines, in expectation of lower future property tax incentives.

## 5.2 Home bias and imperfect capital mobility: direct evidence

### 5.2.1 Incomplete price capitalization

According to proposition 3.2, reducing the subsidy to rental units in location  $j$  is likely to decrease the price of new housing, as demand from investors dries down. The impact of losing eligibility to the subsidy on new unit prices is shown in figure 12, plotting estimates from equation 6, taking the log price of new dwellings sold by developers as the outcome variable. Quantitatively, the effect is however limited: table 5, column (4) shows that prices drop by c. four percent after the end of eligibility. Under standard perfect capital mobility models, required returns to housing investment are given by an exogenous interest rate determined in global financial markets. As a consequence, a place-based housing subsidy to landlords in one location should be accompanied by a full capitalization of the value of the incentive in purchase prices.<sup>69</sup>

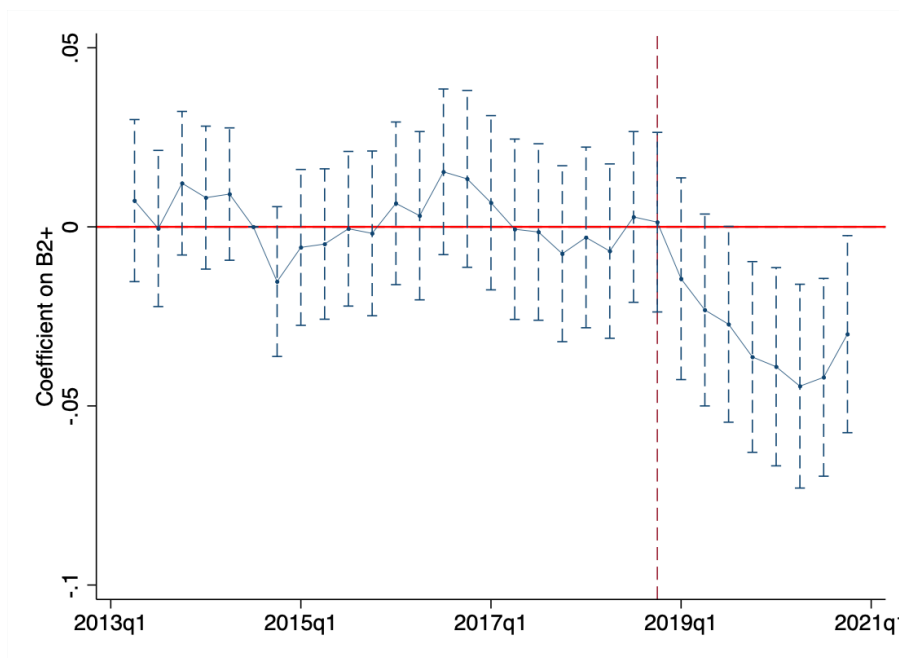


Figure 12: **The impact of exiting the subsidy on new home prices** Figure 12 documents the estimates of coefficient  $\beta_k$  in equation 6 for the impact of losing eligibility to the subsidy (in B2+ cities) on (log) prices of new buy-to-let units by developers, relative to B1 (always-treated) cities. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are clustered at the municipality level, and 95% confidence intervals are shown in brackets. The dashed vertical line marks the end of eligibility of B2+ towns to the *Pinel* incentive relative to B1 towns.

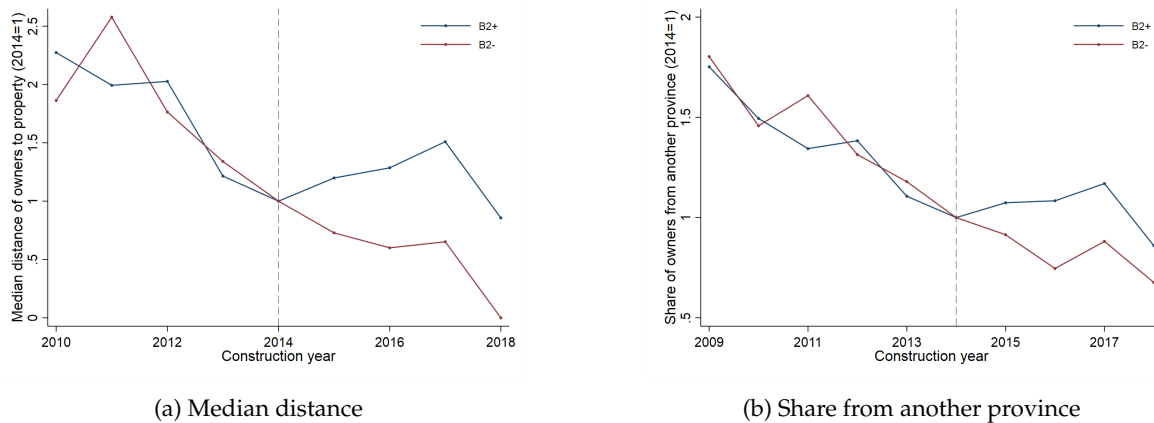
Computations of the value of the subsidy in table 3 (under the null of perfect capital mobility, and for a range of assumptions for the discount rate, baseline rent-price ratio, and bite of affordability requirements) imply a minimum value of 8 to 9 percent (for high discount rates and large potential discounts from rent ceilings relative to market values) and a more central value of 12 to 15 percent of the purchase price (for more standard discount rates and affordability discounts consistent with observed values). However, the observed capitalization effects in new unit prices reach a maximum of about four to five percent after several quarters, only consistent with imperfect capitalization. This provides evidence for a key prediction of proposition 3.2:

<sup>69</sup>See figure 7 and the discussion of proposition 3.2.

to draw additional investment to target areas, place-based policies  $dS_j$  must raise the net return  $r_j^H$  to rental assets in these cities. This implies that part of the incidence falls upon infra-marginal savers. Specifically, because most nearby investors are infra-marginal in their investment choices, due to the presence of home bias, a place-based tax credit benefits local investors, while higher returns are required to make long-distance landlords break even – in terms of full returns, inclusive of their distance costs.

### 5.2.2 Increase in long-distance investment

The spatial investment decision undertaken by landlords weighs the returns of a given residential investment against its (spatially varying) costs - inclusive of any subsidy. As subsidies offset part of the costs, one key prediction of the model is that the reallocation effects on housing investment are stronger among long-distance buyers, who expand the scope of locations in which they consider investing. In essence, long-distance buyers are more likely to be marginal in their decision to purchase rental assets in  $j$ . I test for this impact of buy-to-let investment subsidies on the reduction in home bias. I combine administrative information on the residence location of investors, and the construction year of rental units, for all housing units built over the 2010-2018 period in B2+ (treated) and B2- (control) towns.

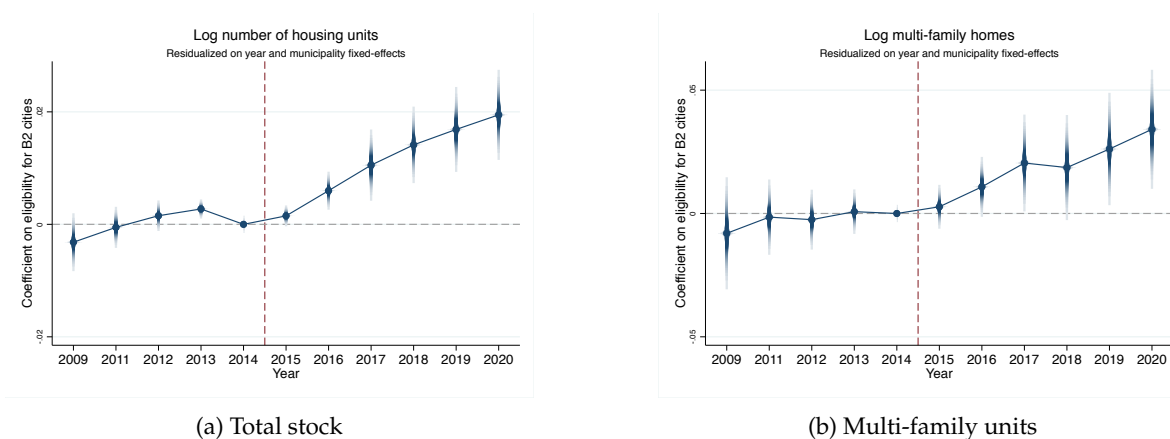


**Figure 13: The spatial ownership effects of entering eligibility** Figure 13 illustrates the spatial reshuffling of housing investment as a consequence of the subsidy. Panel (a) (resp. panel (b)) plots, for B2+ (treated) and B2- (control) towns, the normalized median distance of owners (resp. the normalized share of investors from another province) of rental units in municipalities in each group, depending on the construction year of the rental unit. Statistics are computed from individual ownership data from the 2019 *FIDELI* database.

Figure 13 displays the main result on the impact of place-based subsidies on spatial investment behavior. It compares the residence location of *landlords* for units built in B2+ and B2- locations, depending on the construction year. In particular, I test for the differential evolution, in eligible locations relative to their untreated counterparts, of the median distance of investors from their properties (panel (a)) share of units owned by landlords located in another province (panel (b)). Both outcomes follow similar patterns in the two subsets of B2 municipalities for units built before the divergence in tax credit eligibility, with an increasing home bias and a downwards trend in the presence of remote landlords. However, in B2+ locations that gain eligibility

to the *Pinel* tax credit after 2014, the trend reverses. The increased presence of investors coming from different provinces or living far away in these treated locations is consistent with the model prediction (part (ii) of proposition 3.2). Long-distance investors reallocate their housing portfolio towards cities that represent a low initial share of investments from their origin location, leading on average to a decrease in landlord home bias in locations eligible to the subsidy. This reshuffling of the spatial ownership network is a key mechanism at play in the spatially heterogeneous response of residential investment. By drawing landlords from further away, subsidies partly offset home bias, at the expense of an implicit windfall gain for infra-marginal owners located nearby who would have chosen to invest locally, even absent the subsidy.

### 5.2.3 Heterogeneous effects on the housing stock



**Figure 14: The impact of the subsidy on the housing stock** Figure 14, panel (a) (resp. panel (b)) documents the estimates of coefficient  $\beta_k$  in equation 8 at annual frequency, for the impact of gaining eligibility to the subsidy (in B2+ cities) on the log number of total (resp. multi-family) housing units in the municipality, relative to B2- (never-treated) cities. Data are computed from the *Fichiers Fonciers* exhaustive cadaster. Standard errors are clustered at the municipality level and 95% confidence intervals are shown. The dashed vertical line marks the beginning of eligibility of B2+ towns to the *Pinel* incentive relative to B2- towns.

**Quantity and composition effects on the local housing stock: baseline effects** Using unit-level annual data on all dwellings in each French municipality, I study the impact of the *Pinel* landlord subsidy on the local housing stock in subsidized B2+ locations. Figure 14, panel (a) shows the baseline result, in a fully dynamic difference-in-differences specification: the total stock of housing units in eligible areas increased substantially, by approximately 2 percent in the long-run, after towns gained eligibility to the policy, therefore implying no significant crowding out of un-subsidized units. The make-up of the local housing stock is substantially transformed, with a rise in the share of denser multi-family units, which corresponds more frequently to denser, smaller apartments. Panel (b) shows that the impact of the subsidy is concentrated on multi-family units, with a rise of close to 4 percent in the medium-run; on the contrary, the policy has no effect on the growth of single-family units in treated municipalities. Because the housing market tends to be highly segmented between single-family, mostly owner-occupied housing, and multi-family, generally renter-occupied units, the subsidy to rental investment generates a shift in the composition of the local housing stock, with the apartment share

increasing sharply.<sup>70</sup>

**Heterogeneous effects by openness to outside capital** The third part of proposition 3.2 predicts that, as a consequence of the stronger response of distant investors, the effect of the subsidy on the local housing stock will be stronger in locations more open to outside capital. I estimate the heterogeneous effects of the subsidy on the housing stock, using all B2 locations, according to the following specification:

$$y_{ct} = \alpha_c + \gamma_t + \beta_R \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t \geq \text{Entry start date}} + \theta_R \mathbb{1}_{c \in B2+} \times \mathbb{1}_{t \geq \text{Entry start date}} \times S_c + \Gamma X_{ct} + \epsilon_{ct} \quad (9)$$

where  $S_c$  denotes the dimension of heterogeneity at the level of municipality  $c$ . I define more "capital-open", locations  $c$ , consistent with the model-implied measures described in proposition 3.2, as those situated in a commuting zone  $i$  exhibiting a lower concentration of ownership  $\sum_j \frac{H_{i \rightarrow j}}{H_j} s_{ij}$ , where  $s_{ij}$  is the share of housing assets in  $j$  in the housing portfolio of investors from  $i$ . Since the model also implies that a higher land share in housing production by developers lowers the effective physical supply elasticity of housing, locations with a production technology more reliant on the fixed local supply of land exhibit lower responses. While measuring parameters of the local production technology is challenging, I control for the heterogeneous response of cities by interacting the post-policy indicator with the extent of local buildable land constraints, using the share of urbanized land in the municipality (gathered from satellite information) as a proxy for the bite of land constraints in the spirit of Saiz (2010).

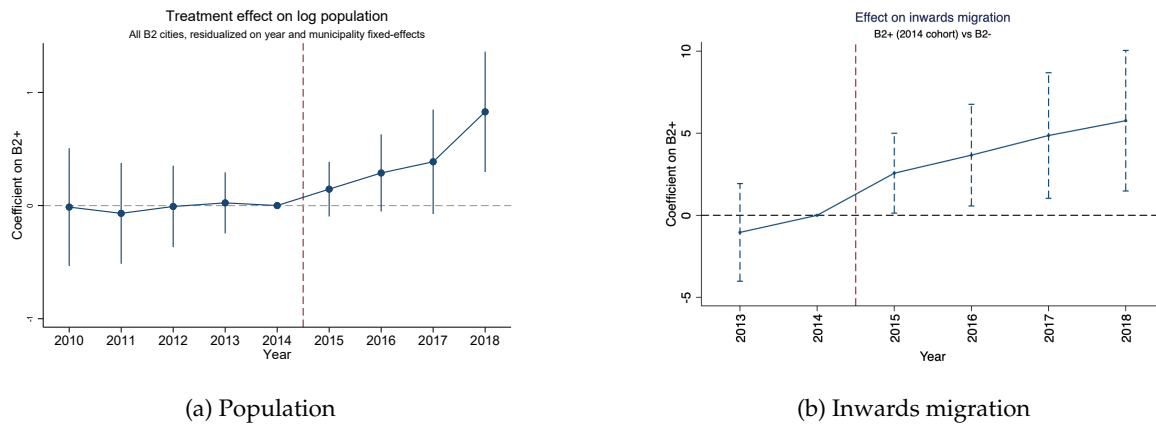
The results, presented in Table 7, provide evidence that the policy witnesses stronger effects in towns that are more open to external capital, as characterized by a smaller CZ-level ownership concentration term  $\sum_j \frac{H_{i \rightarrow j}}{H_j} s_{ij}$ . Across specifications (which include time fixed effects interacted with baseline municipal characteristics, or weighting by population), while the baseline effect of the policy is a substantial medium-run increase in the housing stock of  $c$ . two percent, a higher concentration of ownership in the hands of nearby investors limits the treatment effect. Relying more on local investors indicates that a municipality is more closed to outside capital, and therefore less exposed to the potential reallocation effects of the policy on the portfolio of long-distance owners, a central mechanism through which subsidies raise the local housing stock.

### 5.3 The impact of capital reallocation on worker mobility

Consistent with the theoretical predictions, landlord subsidies increase sales, construction activity, and the local housing stock in targeted areas. A larger local housing stock, however, may lead mostly to an increased per capita consumption of housing, if most of the effect operates through local family formation and household splitting. On the contrary, a last key prediction of the model lies in the fact that the higher local supply of private rental housing induced by tax credits and subsidies affects population growth and inwards mobility. Using town-level data on population and mobility decisions, I examine the impact of the additional private housing induced by the tax subsidy on inwards population movements and overall town population. Panel (a)

<sup>70</sup>Figure D.16 in appendix D.2 evidences an additional induced effect on the local number of social housing units.

of figure 15 graphically presents the results of the estimation specification (8), where the outcome variable is the total population of residents in a town. It plots the coefficients of interest  $\beta_k$  (at annual frequency) corresponding to the differential time-varying impact of being part of the B2+ treated group, relative to untreated B2- cities. In order to focus on medium-term dynamics and because of the low frequency of Census re-sampling, I restrict treated cities to those B2+ towns that had obtained an agreement before the start of the *Pinel* scheme in 2014. The empirical results show that the scheme leads to a slightly more than 1 percent increase in population in treated towns over the long-run, a slightly smaller effect than the housing stock impact.



**Figure 15: The spatial reshuffling of people** Figure 15 illustrates the spatial reshuffling of population as a consequence of the landlord subsidy. Panel (a) (resp. b) plots, for B2+ (treated) relative B2- (control) towns, the coefficient on eligibility in equation 8 for the log population (log in-migrants number) in the town. Standard errors are clustered at the municipality level and 95% confidence intervals are shown in brackets. Statistics are computed from individual Census data.

The composition of the local population also evolves towards lower-income tenants as a consequence of the policy. Using administrative data from the *Caisse d'Allocations Familiales* on social benefits recipients at the town-level, I find an increase in the number of recipients of housing allowances in targeted B2+ locations after the implementation of the policy. Figure viii evidences the differential evolution of the number of recipients of housing vouchers in B2+ and B2- cities over time after 2014, and shows that, while the number of owner-occupier recipients of benefits did not increase, the number of tenants eligible to housing benefits displays a clear rise in B2+ locations relative to B2- controls after the start of the policy in 2014.

Finally, to further investigate the role that spatial mobility plays in triggering additional migration towards treated locations, I show that differential inwards migration trajectories between targeted and untargeted municipalities account for the effect on population. Panel (b) of figure 15 graphically presents the results of the estimation specification (8), taking as the outcome variable the log number of people who moved into a locality within the last year. While data on the origin of residents in the past year is only available starting in 2013, the event-study results demonstrate that, after following parallel trends prior to the policy introduction, B2+ towns treated after 2014 by the landlord subsidy witness a sharp increase in annual flows of inwards migration, drawing additional households and accounting for most of the medium-term rise in population.

Consistent with the model's implied mechanism, shifting the allocation of landlords' portfolios through tax policies induces increased in-migration effects on tenants, and a rise in the population of targeted locations.

## 6 Conclusion

This paper shows that frictions constraining landlord investment behavior across places lead to home-biased landlords, and matter for housing affordability. In a spatial world of footloose renters and immobile landlords, the location of owners themselves shapes the affordability of rental housing. The physical proximity of landlords to a location acts as a force to lower "financial trade costs" in the provision of residential capital and rental services. This agglomeration force operates not through the production (Marshallian knowledge effects) or consumption (scale effects) of the final good, but through financial frictions for long-term capital in a spatially segmented housing market.

Precisely because of spatial barriers specific to residential investment, the geography of *investors* matters for the response of their portfolio allocation to place-based tax incentives, and for the ability of housing supply to accommodate increased demand to live in an area. I estimate empirically the response of residential investment to return shocks, using quasi-experimental variation in a tax credit for affordable new rental housing in targeted French municipalities. The policy not only hastens the sale of existing developer inventory, but also triggers new dwellings construction and purchases, with only limited crowding-out effects on owner-occupiers. The scheme increases mobility towards targeted locations, while shifting the makeup of the local housing stock, and raising the income diversity of eligible communities. Speaking directly to a partial offset of landlord home bias, the subsidy reshuffles the spatial asset ownership network across cities by drawing more remote landlords into the local rental market, and by raising the net return to investment in treated locations.

As housing costs rose steeply in dense urban agglomerations over the last three decades, governments have responded to this trend by intervening in real estate markets, and, in particular, by using tax policy to encourage landlords to provide additional housing in expensive locations. Overall, this paper demonstrates theoretically and empirically that such policies targeting affluent individual landlords can affect the quantity and the allocation of new housing supply. Governments, however, incur large budget costs in the process of implementing these place-based subsidies, because a substantial share of the incidence of these benefits accrues to infra-marginal investors, in particular those whose residence is closest to targeted locations. Whether equity-efficiency trade-offs nonetheless justify locally targeted support to the provision of housing by individual landlords, especially in the presence of heterogeneity in productivity and amenity agglomeration benefits, is a potentially fruitful avenue for future research.

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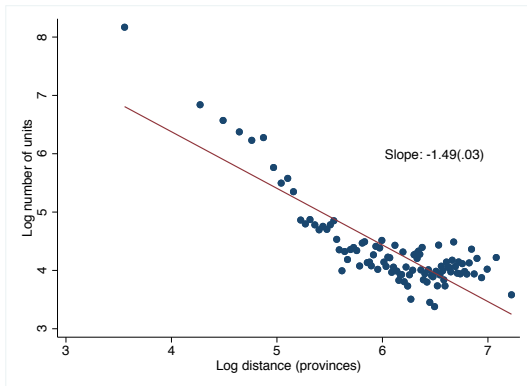
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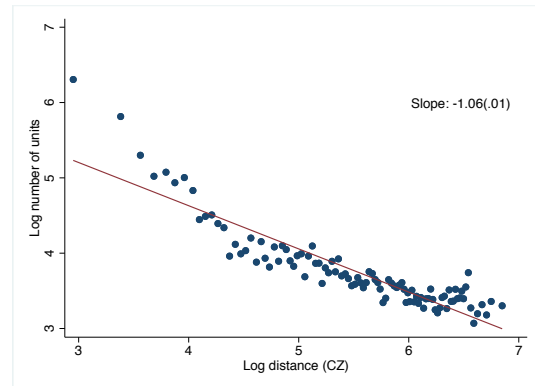
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## Additional main figures

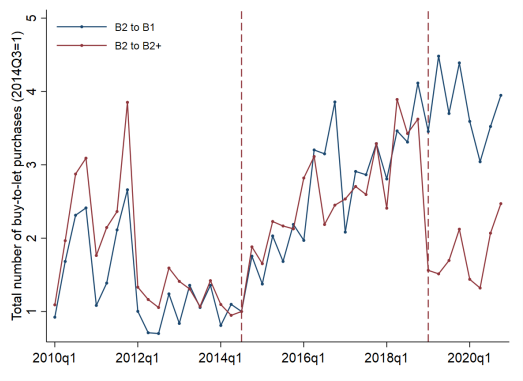


(a) Across provinces

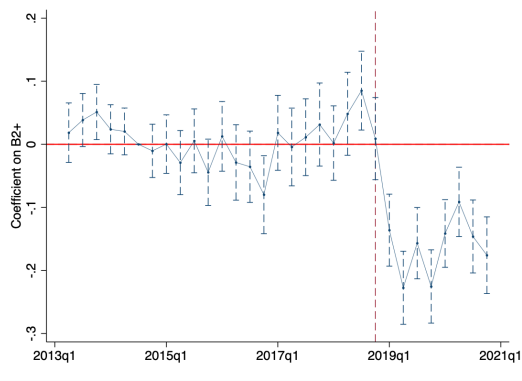


(b) Across commuting zones

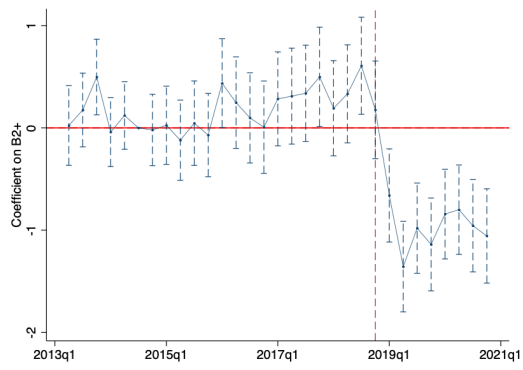
**Figure i: The home bias in homes: additional evidence** Figure i documents additional evidence on the home bias in rental property investment across various aggregation levels. Panel (a) (resp. (b)) plots a binned scatter plot of the log number of units owned in province (resp. CZ)  $j$  by investor from province (resp. CZ)  $i$ . The distance across provinces is the road distance, while the distance across CZ is the haversine distances between their main municipalities. The own-distance is computed as the average distance within a circle with radius equal to the average across CZs or provinces. See table 1 for a description of the data and underlying regressions.



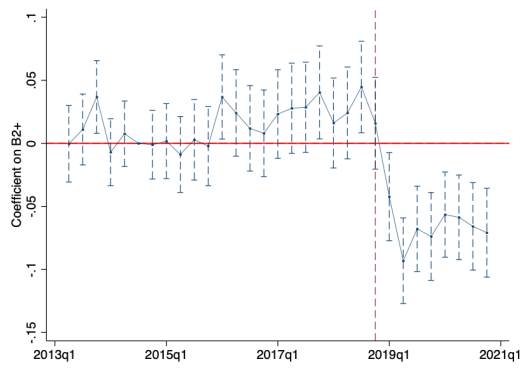
(a) Raw series



(b) Intensive margin: all new home sales

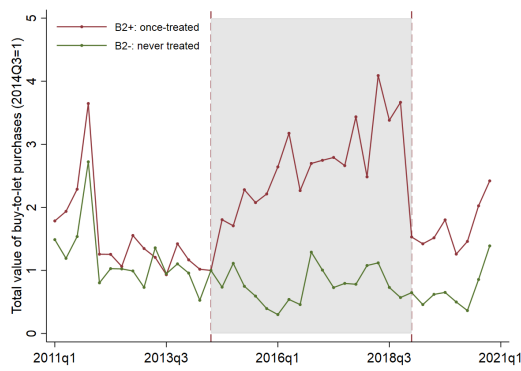


(c) Intensive margin: buy-to-let, log value

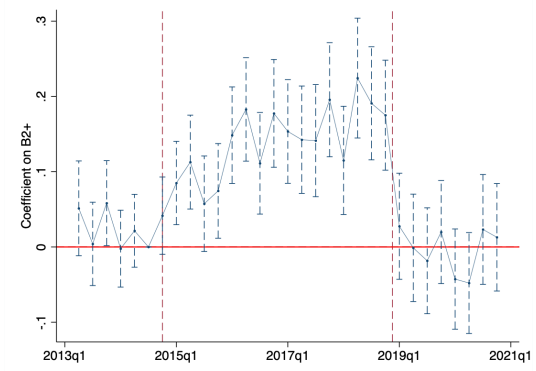


(d) Extensive margin: buy-to-let

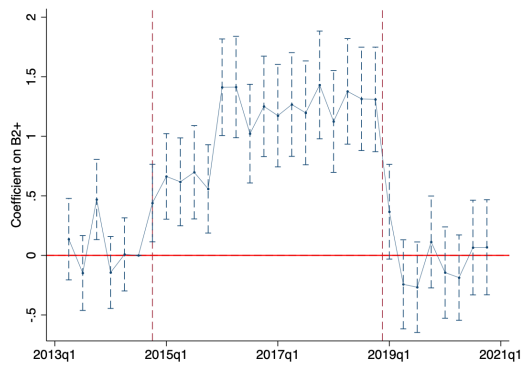
**Figure ii: The impact of exiting the subsidy on new home sales: additional evidence** Figure ii documents additional evidence on the impact of losing eligibility to the subsidy (in B2+ cities) on new home sales by developers, relative to B1 (always-treated) cities. Panel (a) plots the normalized raw series of quarterly new home sales in B2+ and B1 cities among municipalities classified as B2 prior to the September 2014 overhaul. The first dashed vertical line indicate the start of the *Pinel* scheme; the second dashed line marks the end of eligibility for *sales* in B2+ towns. Panels (b), (c), and (d) respectively plot estimates of equation 6 for (log) total new home sales, the (log) value of buy-to-let sales, and the probability of any home sale. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are clustered at the municipality level. The dashed vertical line marks the end of eligibility of B2+ towns to the *Pinel* incentive relative to B1 towns.



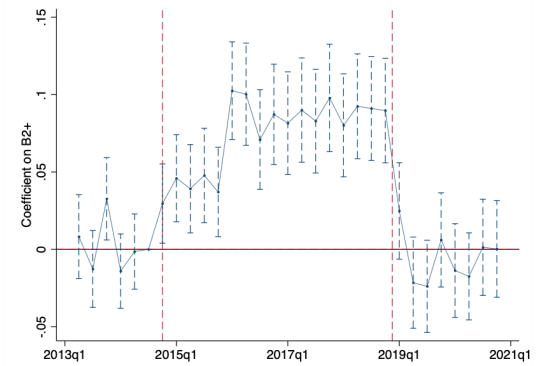
(a) Raw series



(b) Intensive margin: all new home sales



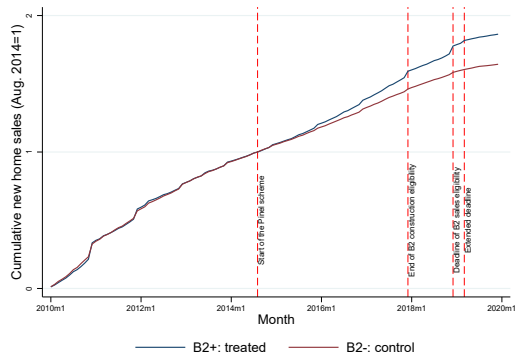
(c) Intensive margin: buy-to-let, log value



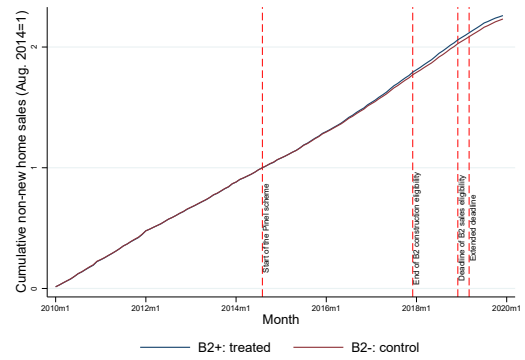
(d) Extensive margin: buy-to-let

**Figure iii: The impact of gaining eligibility to the subsidy: additional evidence** Figure iii documents additional evidence on the impact of gaining eligibility to the subsidy (in B2+ cities) on sales of new homes by developers, relative to B2- (never-treated) cities. Panel (a) plots the normalized raw series of quarterly new home sales in B2+ and B2- cities among municipalities classified as B2 prior to the September 2014 overhaul. Panels (b), (c), and (d) respectively plot estimates of equation 8 for (log) total new home sales, the (log) value of buy-to-let sales, and the probability of any home sale. Standard errors are clustered at the municipality level. The first dashed vertical line indicate the start of the Pinel scheme; the second dashed line marks the end of eligibility for *sales* in B2+ towns. Data are computed from the ECLN database recording individual information on all developer sales.





(a) Sales of new homes

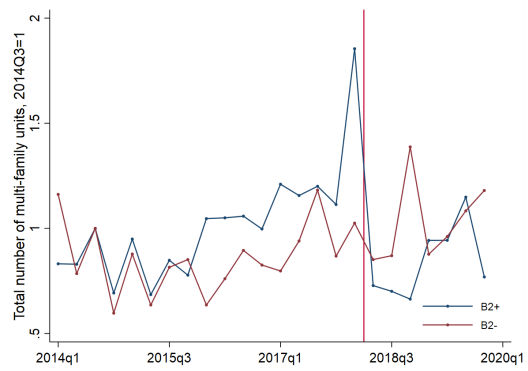


(b) Sales of existing homes

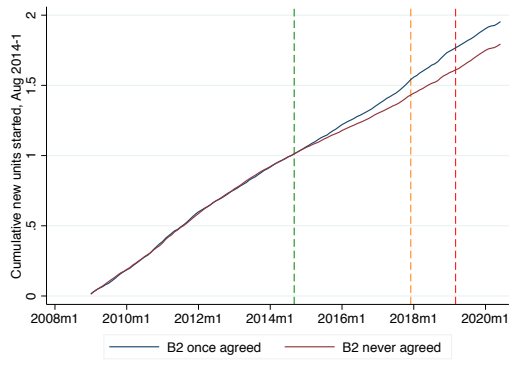
**Figure iv: Placebo test: new and non-new home sales** The figure plots the cumulative value of home sales (normalized prior to the start of the *Pinel* scheme), respectively in never-eligible B2- and once-eligible B2+ municipalities. Panel **a** plots the case of new home sales, which were eligible to the landlord incentive, while panel **b** plots the value of non-new home sales, which were not. Dashed vertical lines mark the beginning, initial planned end, and actual end of the scheme in B2+ towns. Source: *DVF* database, an exhaustive record of housing transactions in France made available by the French Finance Ministry. Sales are restricted to houses and flats with an available transaction value.



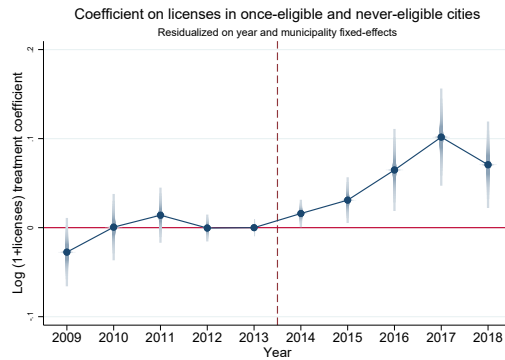
(a) License requests, B2+ v. B1



(b) License requests, B2+ v. B2-



(c) Cumulative housing starts, B2+ v. B2-



(d) Licenses granted: event-study

**Figure v: The impact of gaining eligibility on multi-family housing licenses** Figure v documents additional evidence on the impact of gaining eligibility to the subsidy (in B2+ cities) on the production of new housing, relative to B2- (never-treated) cities. Panels (a) (resp. (b)) plot the quarterly number of licenses *requested* for multi-family units in B2+ cities against B1 (resp. B2-) towns. Panel (c) plots the cumulative monthly number of *housing starts*, respectively in never-eligible B2- (red line) and once-eligible B2+ (blue line) municipalities. The green dashed vertical line indicate the start of eligibility for constructions in B2 municipalities; the orange dashed vertical line marks the end of eligibility for *licenses*, while the dashed red line marks the end of eligibility for *sales*. Panel (d) plots yearly coefficients on eligibility to the *Pinel* incentive for the log number of licenses *granted*, estimating the differential departure from trend in B2+ (relative to B2-) cities. Standard errors are clustered at the municipality level. Sources: *Sit@del2* database, an exhaustive municipality-level repository of all licenses granted - see appendix A for a detailed description of the data.

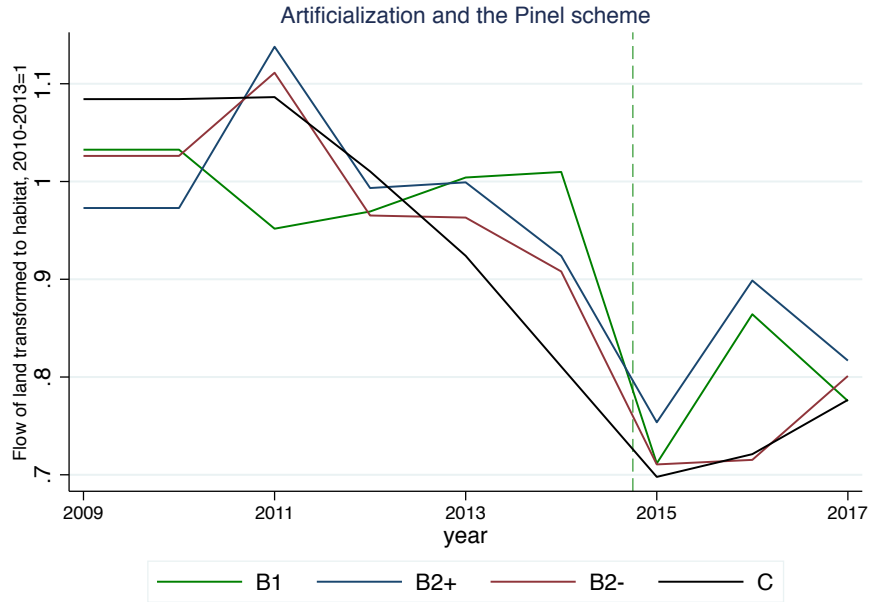


Figure vi: **Total land area shifted from natural to residential uses** The figure plots the annual area of land converted from natural to residential use in France from 2009 to 2017, respectively in B1 (green line), once-eligible B2 (blue line), never-eligible B2 (red line), and never-eligible C (green line) municipalities. The annual flow of land converted is normalized to the average of the period immediately preceding the implementation of the policy (2009-2014). Sources: *Observatoire de l'artificialisation des sols* database - see appendix A for a detailed description of the data.

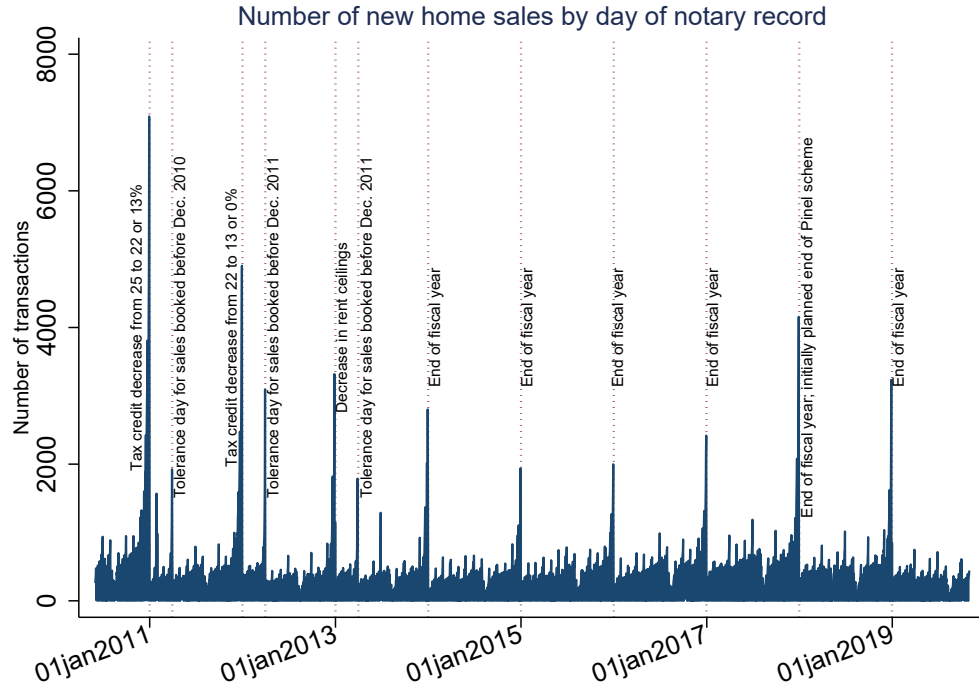


Figure vii: **New home sales recorded by notaries, weekly** The figure shows the total weekly number of new housing units sold in France from 2010 to 2019. Dashed lines indicate eligibility deadlines (or deadline extensions) before the switch to a less generous tax regime, as well as the end of each fiscal year. The successive tax regimes are described in table 4 and appendix B.1. The total number of units sold is computed from microdata in the *DV3F* database, an exhaustive registry of all housing deeds in the country. The “new” unit status is inferred from the VAT treatment of the dwelling.

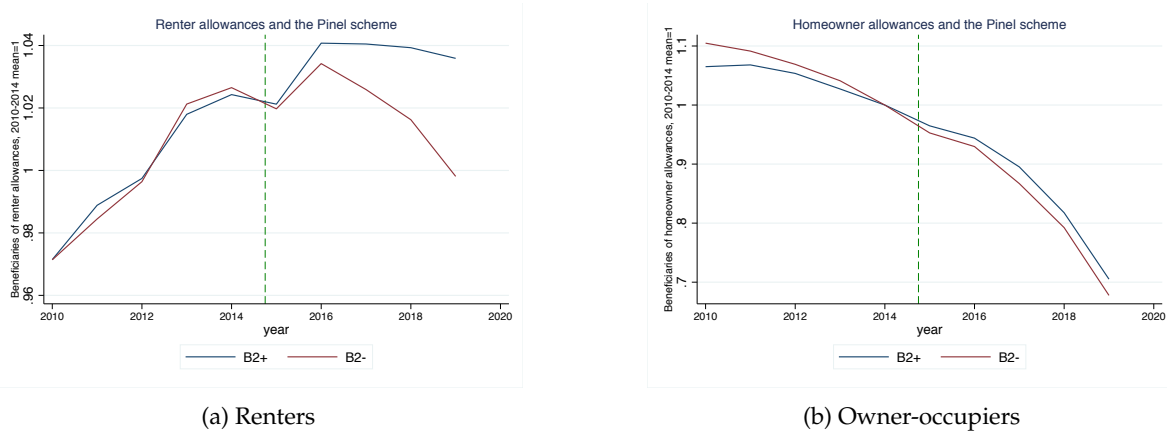


Figure viii: **The evolution of recipients of social benefits** Figure viii illustrates the evolution of the local population as a consequence of the landlord subsidy. Panel (a) (resp. b) plots, for B2+ (treated) relative B2- (control) towns, the normalized number of recipients of social benefit allowances for rental housing occupiers (resp. owner-occupiers) in the town. Statistics are computed from *Caisses d’Allocations Familiales* municipality-level data.

## Main tables

Table 1: The home bias in homes: gravity estimates

Across provinces					
Distance coefficient	-1.206****	-1.489****	-1.241****	-0.139****	-0.108****
	(0.0277)	(0.0120)	(0.0109)	(0.0116)	(0.0109)
Two-way origin-destination FE	No	Yes	Yes	Yes	Yes
Includes <i>ii</i> pair	Yes	Yes	No	No	Yes
SCI controls	No	No	No	Yes	Yes
R-Square	0.303	0.815	0.843	0.944	0.954
Observations	7722	7721	7625	7625	7721
Across commuting zones					
Distance coefficient	-0.573****	-1.063****	-0.779****	-1.158****	
	(0.0107)	(0.0111)	(0.0519)	(0.00709)	
Two-way origin-destination FE	No	Yes	Yes	Yes	
Includes <i>ii</i> pair	Yes	Yes	Yes	No	
Method	OLS	TWFE	PPML	TWFE	
R-Square	0.138	0.440	0.2446	0.699	
Observations	17976	17959	17959	17655	

Standard errors in parentheses. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$

Table 1 estimates the impact of bilateral distance on the log number of units  $C_{ij}$  owned by investors from location  $i$  in location  $j$ . It reports the coefficient  $\beta$  on (log) distance in two-way fixed effects regressions of the form  $\log C_{ij} = \alpha_i + \gamma_j + \beta \log d_{ij} + \eta X_{ij} + \epsilon_{ij}$  where  $X_{ij}$  is a (potentially empty) control for the social connectedness index (Bailey et al., 2018) between two locations. The top panel plots the estimated effect across provinces, while the bottom panel estimates the effect of log distance on log ownership across commuting zones in mainland France. To comply with disclosure confidentiality requirements, I aggregate rental dwelling ownership at either the commuting zone level (across CZ or "zones d'emploi") in mainland France or at the province level (across provinces or "departements" in mainland France), and exclude bilateral pairs of origin-destination locations with ten units or fewer.

Table 2: Summary statistics: municipalities

	Across classifications					
	Abis	A	B1	B2+	B2-	C
<i>Town population</i>	66,520	16,747	8,849	8,320	1,830	787
<i>Median income</i>	26,722	24,417	23,587	21,197	22,334	19,935
<i>% in private rental housing</i>	25.6	17.4	17.1	18.5	14.4	14.8
<i>% moved in less than 2 years ago</i>	12.0	11.1	11.4	11.6	9.1	8.7
<i>% in multi-family units</i>	80.6	40.1	26.2	26.8	11.1	6.2
<i>Monthly rent (EUR/sqm)</i>	20.8	15.6	12.3	10.3	10.8	8.4
<i>House price (EUR/sqm)</i>	5,487	3,596	2,692	1,935	1,971	1,305
<i>Real estate sales (2014-19)</i>	6,387	1,631	875	784	121	47
<i>Municipalities</i>	77	650	1,417	905	2995	30,507

Table 2 presents within-zone simple averages of socio-demographics characteristics for municipalities across the *Pinel* classification categories described in section 4. The data for average population, median income, the share of households in private rental housing, the share moving less than 2 years ago, and the share of the population in multi-family units, are from 2014, at the start of the *Pinel* scheme. The monthly rents are from the cross-section of 2017-2018 provided by the French Housing Secretariat. The house prices and average number of transactions are from the DV3F database over the 2014-2019 period.

Table 3: Net present value of the *Pinel* incentive

		Discount rate			
		5%	3%	2%	1%
<b>Depth of rent markdown</b>	<b>5%</b>	12.4%	13.6%	14.3%	15%
	<b>10%</b>	10.7%	11.7%	12.2%	12.8%
	<b>15%</b>	8.9%	9.7%	10.2%	10.7%

The net present value of the incentive embedded in the *Pinel* scheme is computed by assuming that the landlord receives the full benefit of the tax credit (2 percent of the purchase price each year), a 5% baseline rent-to-price ratio  $\lambda$ , and a nine-year commitment to affordable rents. The value is then given by  $\sum_{k=1}^9 \frac{0.02-\lambda d}{(1+r)^k}$ , where  $d$  denotes the markdown of controlled rents relative to baseline market rents, and  $r$  is the assumed discount rate. See appendix B.1 for a discussion of the assumptions.

Name	Period	Mechanism	Level	Affordability requirements	Increased/decreased generosity
<i>Perissol</i>	Jan. 1996-Jul.1999	Depreciation	80% (Y1-Y4: 10%-Y5-Y24: 2%)	None	Increased
<i>Besson</i>	Aug. 1999-Dec.2002	Depreciation	50% (Y1-Y5: 8%-Y6-Y9: 2.5%) option for Y10-Y15:2.5%)	Rent ceilings Resident income ceilings	Decreased
<i>Robien</i>	Jan. 2003-Aug.2006 Sep. 2006-Dec.2009	Depreciation Depreciation	50% (Y1-Y5: 8%-Y6-Y9: 2.5%) (option for Y10-Y15:2.5%) 50% (Y1-Y7: 6%-Y8-Y9: 4%)	Rent ceilings (close to market rate) No resident income ceilings Reduced rent ceilings	Increased Decreased
<i>Borloo</i>	Sep. 2006-Dec.2009	Depreciation + tax deduction	50% (Y1-Y7: 6%-Y8-Y9: 4%) +30% deduction off rental income	Rent ceilings (20% below Robien) +Resident income ceilings	Decreased
<i>Scellier</i>	Jan. 2009-Dec.2010 Jan. 2011-Dec.2011 Jan. 2012-Dec.2012	Tax credit	25% 22% (13% if not energy efficient) 13% (6% if not energy efficient)	Rent ceilings (close to market rate) No resident income ceilings 9 year commitment	Increased Decreased Decreased
<i>"Social" Scellier</i>	Jan. 2009-Dec.2010 Jan. 2011-Dec.2011 Jan. 2012-Dec.2012	Tax credit + tax deduction	Same as Scellier each year +5% per 3-year additional period +30% deduction off rental income	Rent ceilings (20% below Scellier) Resident income ceilings 9 year commitment (12-15 year option)	Increased Decreased Decreased
<i>Duflot</i>	Jan. 2013-Aug. 2014	Tax credit	18% (Y1-Y9: 2%) 12-21% (Y1-Y6: 2%) (Y7-Y9: 2%) (Y10-Y12: 1%)	Rent & resident income ceilings based on "Social Scellier" levels	Unchanged/decreased
<i>Pinel</i>	Sep. 2014-Dec. 2021	Tax credit		Higher rent ceilings Higher resident income ceilings	Increased

Table 4: Incentive schemes for buy-to-let new residential housing in France (1996-2021)

Table 5: Eligibility exit: estimates

	(1)	(2)	(3)	(4)
Dependent Variable	New home sales	Buy-to-let sales (int. margin)	Buy to let (ext. margin)	Prices
3 quarters before exit	-.0428062 (.0405902)	.0019091 (.0301027)	.016115 (.0182122)	-.0029532 (.0128662)
2 quarters before exit	.010333 (.0435246)	.0483503 (.0335806)	.0240921 (.0185643)	-.0067724 (.0124153)
1 quarters before exit	.0660347*** (.0415818)	.0851022*** (.0319081)	.0447099*** (.0185127)	.0027679 (.0121608)
1 quarter after exit	-.1660149*** (.0412847)	-.1361208*** (0.0291161)	-.0423949*** (.0178303)	-.0144896* (.0143686)
2 quarters after exit	-.2383592*** (.0417033)	-.2275238*** (.0295576)	-.09317*** (.0173303)	-.0232125** (.0136663)
3 quarters after exit	-.1778878*** (.040539)	-.1567859*** (.028835)	-.0678629*** (.01725)	-.0272029** (.0139467)
4 quarters after exit	-.218649*** (.0408973)	-.2254706*** (.0296424)	-.0738466*** (.0177864)	-.0363281*** (.0135863)
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	90,687	90,687	90,687	32,438

Table 5 estimates the impact of losing eligibility to the incentive scheme for B2+ municipalities (relative to B1) after 2018Q4. Columns 1, 2, 3, 4 respectively plot estimates of equation 6 for (log) total new home sales, the (log) number of buy-to-let sales, the probability of any home sale, and the price of new units. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are in parentheses clustered at the municipality level. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$



Table 6: Eligibility exit: estimates

Dependent Variable	(1) New home sales	(2) Buy-to-let sales (int. margin)	(3) Buy to let (ext. margin)
2 quarters before entry	-.0022547 (.0260417)	.0059869 (.0164589)	-.0140338 (.0122359)
1 quarters before entry	.0215091 (.0247059)	.0162246 (.0166583)	-.001393 (.0123971)
1 quarter after entry	.0415695*** (.0262121)	.0774417*** (.0196763)	.0296224*** (.0130586)
2 quarters after entry	.0849734*** (.0282319)	.1065969*** (.0210541)	.0459934*** (.0143595)
3 quarters after entry	.11272*** (.0318185)	.1266174*** (.0231026)	.0392008*** (.0145558)
4 quarters after entry	.0574344** (.032381)	.1202703*** (.0234635)	.047749*** (.0155446)
15 quarters after entry	.2243742*** (.0406162)	.254235*** (.0300801)	.0924642*** (.017272)
Municipality FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Observations	134,208	134,208	134,208

Table 6 estimates the impact of gaining eligibility to the incentive scheme for B2+ municipalities (relative to B2-) after 2014Q3. Columns 1, 2, and 3 respectively plot estimates of equation 8 for (log) total new home sales, the (log) number of buy-to-let sales, and the probability of any home sale. Data are computed from the ECLN database, recording exhaustive dwelling-level information on all developer sales. Standard errors are in parentheses clustered at the municipality level. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$

Table 7: Effects on the local housing stock

<b>Baseline effects</b>				
"Eligibility x Post"	.0116819*** (.0025491)	.0159897*** (.0026204)	.0096321** (.004432)	.0233584*** (.0063502)
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Time-varying controls	No	Yes	Yes	Yes
Weighting by pop.	No	No	Yes	Yes
HS elasticity	No	No	No	Yes
Observations	42,631	40,598	40,598	40,349
<b>Heterogeneous effects</b>				
"Eligibility x Post"	.037374*** (.0113679)	.0423141*** (.0110031)	.0222928 (.0148743)	.0346503** (.0147433)
"Eligibility x Post x $S_c$ "	-.0507972** (.0222711)	-.0522198** (.0213539)	-.0250971 (.0358387)	-.0226713 (.0361204)
Municipality FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Time-varying controls	No	Yes	Yes	Yes
Weighting by pop.	No	No	Yes	Yes
HS elasticity	No	No	No	Yes
Observations	42,094	40,114	40,114	40,114

Regressions coefficients are shown for the estimation of equation 9 in the sample of B2 municipalities. The "Eligibility x Post" coefficient indicates the DiD estimate of  $\beta_R$ , the interaction of post-entry date and eligibility (membership of the B2+ group). The "Eligibility x Post x  $S_c$ " coefficient is the estimate of  $\theta_R$ , the heterogeneous impact of eligibility on the local housing stock by the extent of  $S_c = \sum_i \frac{H_{i \rightarrow j(c)}}{H_{j(c)}} s_{ij(c)}$ , the model-implied measure of the concentration of ownership for the commuting zone  $j$  to which  $c$  belongs. The model predicts that a more dispersed ownership (a lower  $S_c$ ) entails a stronger effect of subsidies on housing production. "Time-varying controls" include the interaction of year fixed-effects with initial characteristics of the municipality (median income and population). "Land share" indicates the presence of controls for the interaction of "Eligibility x Post" with the baseline urbanized land share in the municipality. Standard errors are in parentheses clustered at the municipality level. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$ .

# Appendices

## A Data construction

I use data drawn from a wide array of administrative sources, including exhaustive individual data on the universe of building licenses requests, housing starts, land purchases, housing transactions, developer residential sales, and social housing units in France. A confidential and exhaustive panel of all taxpaying households in France allows me to track the individual determinants and consequences of investors' decisions to invest in the housing sector. Finally, I also rely on municipality-level measures of aggregate outcomes, spatial and housing mobility, as well as municipality-industry measures of employment, employee earnings, and firm creation for all c. 35,000 municipalities in France, in order to assess the consequences of place-targeted housing policies at the local level. This appendix provides additional detailed information on the construction of the data, sample selection, and data transformation performed.

### A.1 Housing markets

**Home sales** To study the impact of the policy on local markets for both newly built units and existing homes, and track its role for local home prices, housing market activity and land values, I use two main sources of administrative data on home sales. First, I avail myself of exhaustive and never-used before data recording all housing and land transactions in France from January 2010 to December 2019, recorded by the French Treasury for tax purposes in the *DV3F* database. For home sales, I restrict the sample to sales of entire houses and apartments, which represent around 850,000 annual housing transactions over the sample period. Transaction details include the name of the municipality, the exact geo-coding of the address, and, for most transactions, the price and exact square footage of the unit sold. Using specific legal codes defining the VAT and transfer tax treatment of the sale, I am able to determine whether a sale corresponds to a new or an existing home, and whether the sale was an "off-plan" sale (*Vente en l'Etat Futur d'Achevement* or VEFA) - the main mechanism through which developers sell rental property in France.

Second, I use a restricted-access exhaustive survey of developers, the *Enquete sur la Commercialisation des Logements Neufs* (ECLN), which precisely records detailed information on all new housing sales performed by developers at the unit level. Relative to the DV3F data, the ECLN includes additional information on the characteristics of the unit, the identity of the developer, and, for most units since the 2010-2011 wave, the final destination of the purchase (owner-occupied housing, subsidized or unsubsidized rental property, social housing...), as well as its exact location. The survey is exhaustive and of high quality, with projects being followed over time for several quarters; it is however restricted to housing projects comprised of five or more units, thus including all large multi-family projects but potentially missing small individual constructions still eligible under the Pinel scheme.

**Land transactions** I avail myself of two main sources for land sales and prices in France. First, I exploit the *DV3F* database but restrict transactions to sales of constructible land ("*terrains a batir*"). Second, I use the

*Enquete sur le prix des terrains a batir*, an exhaustive survey recording the purchase price, final use, and overall area of all constructible land sold in France every year, to examine the consequence of the policy for land prices.

**New construction** To evaluate the impact of the *Pinel* scheme on house-building, I use administrative confidential data from the *Sit@del2* database combined by the French Housing authority. The *Sit@del2* data contain exhaustive information on the universe of housing permits requested and include details about the exact date of the request, its approval date and current status, the geographic location, a summary description of the project, and the number and type of units built, for each housing project. The database allows me to aggregate all housing authorizations and housing starts (separately for single- and multi-family units), at the dwelling level across all municipalities in France.

**Housing stock** I rely on exhaustive annual data from the cadaster - the so-called *Fichiers Fonciers*, a tax-based registry of the universe of housing units in France, which records extensive information on all housing units at yearly frequency, as of January 1<sup>st</sup>. I aggregate data at the municipal level, in order to compare a variety of relevant outcomes between eligible and uneligible locations, including the total number of local dwellings and the share of rental and multi-family units. As a complement, I also use data from INSEE, the French statistical institute, which combines the annual Census with other sources (including INSEE's regular Housing survey, housing tax files, and other information) to produce information on all housing units – occupied or not – in France. The resulting *Base Logements*, available at yearly frequency until 2017, provides information on the quantity, use, and nature of all housing units in a municipality, including detailed information on tenure status, available amenities, turnover rates, and ownership status.

**Social housing** To study the extent of crowding-in or crowding-out of social housing by intermediate housing construction incentivized through the *Pinel* scheme, I use the *Répertoire des logements locatifs des bailleurs sociaux*, an exhaustive and legally mandated repository of all social housing units in France. The RPLS provides an annual assessment of the number of social housing units (occupied or vacant) in the municipality, starting in 2012. Since reporting to the RPLS is mandatory for all owners of social housing, its data provides a highly reliable source of information on the extent, exact location, and characteristics of new social housing supply, at yearly frequency, for all municipalities in France.

## A.2 Investor behavior

**Tax returns** I use the universe of French individual tax returns from 2006 to 2019, an exhaustive sample of more than 30 million tax units each year (collectively known as *Fichiers POTE - panelisables*). The data, the panel structure of which allows me to track taxable households over time, include detailed information on household characteristics, all forms of taxable capital, pension, and labor income, any tax deductions and shelters used, the amount (if any) invested under subsidized housing schemes. In order to measure landlord status, I define as landlords all households who file a so-called *2044* schedule, and register either a strictly

positive amount of income from rental housing, or deduct a strictly positive amount of rental losses (“negative gearing”). I merge individual tax returns data to mandatory additional schedules filled by initial investors in the subsidized rental investment scheme (known as schedule *2044-EB*, for *engagement bailleur*, i.e. “landlord commitment”): for almost all buyers under the *Pinel* scheme who fill an online tax return in 2016, these data record the exact location of the property they purchased, the price paid, the square footage of the unit, as well as additional information on the tenant and unit characteristics.

**Housing stock ownership** I make use of the FILOCOM/FIDELI database, an exhaustive repository of all c. 37 million housing units (as of 2019) in France, available every other year (and every year starting after 2018). Data on individual housing units obtained from the property tax collection process are merged by tax authorities to summary information from income tax returns on resident households. In addition to several pieces of information on the dwelling, including the income and household composition of its residents, its construction year, or the last year in which it was sold or donated, the database records the municipality of the address to which the property tax (the *taxe foncière*) is sent. For each non-owner-occupied property, this allows me to observe the residence location of the rental investor for each rental dwelling, a key input in my gravity approach to housing investment. I match these data to a matrix of great-circle, road, and cultural distance between the commuting zones or provinces to which the municipalities of the owner and the renter belong, in order to examine the role of spatial frictions in residential investment behavior.

**Wealth and assets** Since tax returns data do not record exhaustive information on overall assets and wealth (see e.g. Garbinti, Goupille-Lebret, and Piketty (2021) for an attempt at reconciling tax and survey data on wealth in France), I also exploit the *Enquete Patrimoine (EP)*, a French equivalent to the U.S. Survey of Consumer Finances, which records detailed information on household income, assets, and portfolio composition. The survey is a short panel which interviews household twice over a four-year period (I use the 2018 vintage). It is matched to summary information on income and family status from household tax returns in the past year, and over-samples wealthier tax units. I use restricted-access microdata from the *EP* in order to measure landlordship and its correlates at the individual level. The data allow me to observe, for each surveyed individual, the level and allocation of their wealth across various asset types, including the value and nature of housing assets owned by a household. Landlords are defined as individuals who own at least one housing unit (in addition to their primary residence) that they lease for the entire year: these correspond to c. 10 percent of French households (c. 3 million).

### A.3 Data on the local economy

**Population** I use data on municipal population assembled by INSEE through the annual Census, to estimate the extent of population mobility induced by the scheme towards targeted areas. In addition, I employ aggregate data on municipality-to-municipality annual migration flows from the French statistical institute INSEE, to measure the extent to which the policy induced population mobility towards targeted areas, and away from

untargeted regions. To study the composition of aggregate flows and assess how the socio-demographic composition of renters and movers, I use individual data from the Census on mobility decisions from INSEE's *Fichiers details - Mobilité résidentielle*.

**Renters and low-income households** To look at specific mobility effects among lower-income households who rent, I rely on data from the *Caisse d'Allocations Familiales*, France's main outlet for social benefit payments. In particular, the CAF records the annual number of beneficiaries of each of its three main housing allowances, the APL (*Allocation Personnalisée au Logement*), ALS (*Allocation de Logement Sociale*), and ALF (*Allocation de Logement Familiale*), in each municipality, through the *Fileas BCA* database. These allowances are designed to help mostly renters, but also homeowners, cover their regular housing expenditures.<sup>71</sup> I use the *Fileas BCA* data to obtain counts of beneficiaries of any housing allowance, as well as separate counts of renters and homeowner beneficiaries. I also use the CAF files to measure the percentage of poor households in the local population, defined as those earning less than sixty percent of median income per consumption unit.

**Local economic activity** To test whether the policy affects local economic activity, I use a series of administrative databases on municipality-level outcomes. I first exploit establishment-level data from the *Repertoire des Entreprises et des Etablissements* database,<sup>72</sup> comprised of information on the number of employees, detailed industry classification, and exact location for each establishment in France, which I aggregate at the municipality-year level, to estimate the annual stock of firms, as well as entry and exit rates, in each industry and municipality, before and after the implementation of the policy. I also use data from the French Social Security system (*Unions de Recouvrement des cotisations de Sécurité Sociale et d'Allocations Familiales* or URSSAF), aggregated at the municipality-industry-year level, on the total number and overall wages of salaried workers by 4-digits industry, to study the impact of the policy on employment and wage rates in the real estate and construction sectors.

**Land use** The impact of subsidizing housing supply on urban sprawl and expansion requires granular, geolocalized information on land use and land cover. I avail myself of the *Observatoire de l'artificialisation des sols* database, which combines satellite information with administrative records, and constitutes the most precise and exhaustive repository of information on land use at a highly granular definition (at the level of "parcelles", or cadastral plots). The data allow me to compute the amount of land converted each year from "natural" use to residential or industrial use, within each French municipality. I also use the *Corine Land Cover* database, a satellite measure of land use, to assess the share of urbanized land in each town.

#### A.4 Institutional arrangements

**Subsidized housing scheme eligibility** I exploit data from the French Housing Ministry detailing the eligibility status and zoning rules applicable to each municipality in France. The data record the zone applicable

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<sup>71</sup>Bozio et al. (2015) provide a detailed overview of France's various personal housing benefits.

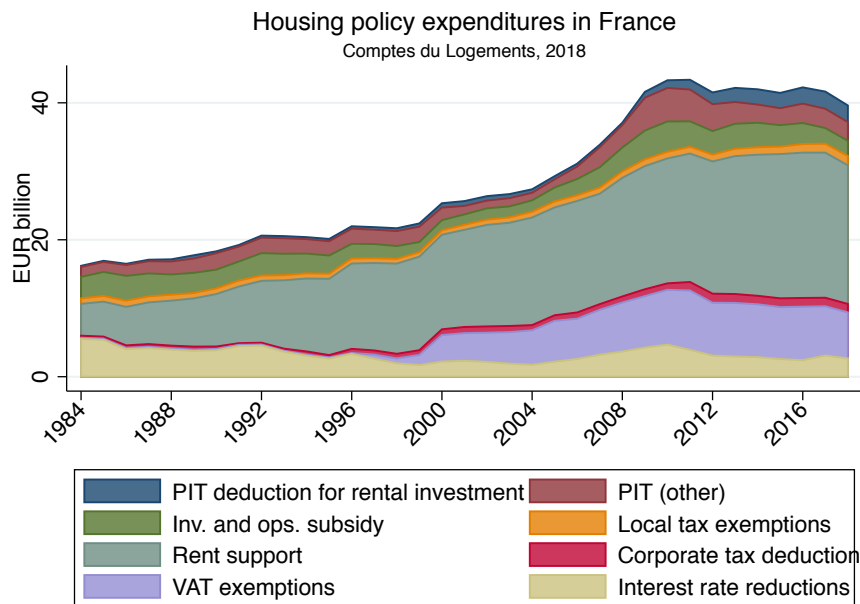
<sup>72</sup>Base SIRENE (BPE) - 2007-2019, INSEE [producteur], ADISP [diffuseur].

to a municipality ("A", "Abis", "B1", "B2", or "C"), both from 2009 to the September 2014 change and afterwards. The law determines the eligibility of an area to the subsidized investment scheme, and, for eligible locations, both the rent ceiling per square meter applicable and the maximum income of the residents. I also gathered data on all individual eligibility agreements granted to B2 or C municipalities under the Pinel scheme. I obtained restricted-access data from the Ministry of Defence on the coverage of towns by "garrison towns" revitalization contractual agreements (*Contrats de Revitalisation des Sites de Defence*), since the 2018 reform expanded Pinel eligibility to cover towns that had lost a military unit and were covered by such a revitalization contract (see B.3). Finally, to study the diverging dynamics of neighbouring towns when they differ by eligibility status, I also employ an adjacency matrix of all municipalities provided by OpenStreetMap, an open-source software recording adjacency links in the graph of all c. 35,000 French municipalities.

## B Institutional background

### B.1 Subsidy schemes to new housing in France

**Rental property policies in France** France spends substantial amounts of budget support for housing policies. As of 2019, the country dedicated c. €40 billion a year (around \$45 billion, close to two percent of French GDP) to various housing market demand and supply-side subsidies. These include around EUR 22 billion in direct rental assistance to low- and middle-income tenants, c. EUR 3 billion in interest rate subsidies for social housing construction and exploitation, and interest-free loans to middle-income owners; c. EUR 2 billion in building subsidies, and around EUR 13 billion in a variety of fiscal expenditures, from reduced corporate tax and VAT rates for social housing construction, to tax deductions for environmental upgrades of primary residences, to tax credits associated with rental property investment (see figure B.1).



**Figure B.1: French housing policy costs** Various housing policy expenditures in France were estimated to cost close to 40 billion € as of 2018, or c. 2 percent of French GDP. They include personal income tax deductions for rental investment, the focus of this paper, as well as a variety of other tax expenditures, VAT/corporate tax/local taxation exemptions, interest rate subsidies, operations and investment subsidies to housing developers, or direct cash allowances to tenants. Figures are from *Comptes du Logement*, an annual report compiled by the French Housing Secretariat on the total costs of policies supporting housing. Sources: *Service de la donnée et des études statistiques, Sous-direction des Statistiques du Logement et de la Construction*.

**Tax credits for rental property investment** The latter are the focus of the empirical exercise in this paper. They allow for either accelerated depreciation deductions off the investor's personal income tax basis,<sup>73</sup> or direct reductions in the investor's tax liability. A succession of tax schemes were put in place starting in the 1990s, with varying deduction rates and limitations related to affordability guidance, targeted areas, rent ceil-

<sup>73</sup>In France, depreciation is generally not allowed as a deductible expense for owners of rental property, unless the unit is furnished and rented under a specific tax status, the *Loueur en meuble non professionnel* for a short period of time.



ings, and tenant income conditions. They were renamed after each Housing Minister who introduced changes. Table 4 provides summary information on the schemes, which are describe in more detail below.<sup>74</sup>

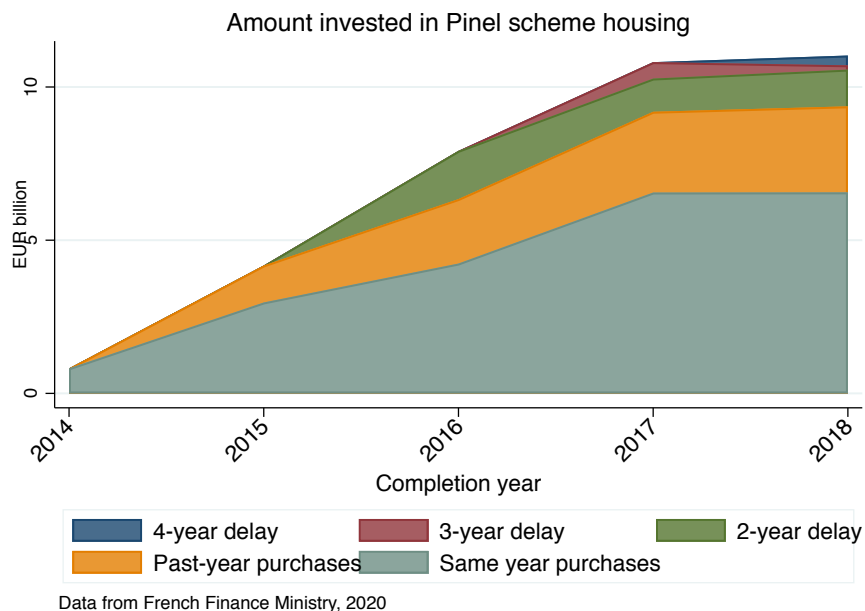
- From 1996 to August 1999, the *Perissol* scheme was in place. It allowed households to depreciate 80 percent of the purchase cost of the rental property (10 percent of the investment in the initial four years, then another 2 percent a year for 20 years, with no affordability requirement). The depreciation could be used to reduce taxable rental income, and any excess could offset overall taxable income up to a limit of c. EUR 15,000 per year.
- From 1999 to 2002, the *Besson* law reduced the generosity of the accelerated depreciation mechanism. It allowed a depreciation of only 50 percent of the cost of newly built residential assets to offset the individual income tax liability of the investor over nine years (8 percent in the first five years, then 2.5 percent a year for four years), with an option for an additional 2.5 percent a year depreciation for 6 years, pushing the total potential for tax depreciation to 65 percent over 15 years. The cap on excess depreciation (negative gearing against non-rental taxable income) was lowered to c. EUR 10,000 annually. Moreover, below-market rent ceilings and maximum resident income conditions were put in place, and made dependent on whether the investment occurred in Paris, in large cities, or elsewhere, with a view to turn the scheme into a form of inclusionary zoning.
- From 2003 to 2006, the more generous *Robien* law kept the same depreciation conditions as the *Besson* law, but raised rent ceilings by 10 to 50 percent (depending on the location) and abolished the maximum resident income provision, therefore increasing net incentives for taxable households to invest in individually-owned rental housing.
- From 2006 to 2009, the so-called “re-focused” *Robien* law slowed down the pace of allowed depreciation (to 6 percent a year for seven years, and four percent for an additional two years), removed the option to extend it for another six years, and lowered rent ceilings in low-density locations. An alternative scheme, the *Borloo* law, allowed households to add to the basic *Robien* tax shelter an option for a 30 percent deduction off rental income, under the condition that rents be capped c. 20 percent below *Robien* rent ceilings.
- From 2009 to 2010, with a view to kickstart the recovery after the Great Financial Crisis, the newly introduced *Scellier* scheme reshuffled tax subsidies to owner-operated rental housing. It changed the definition of eligible areas and excluded small towns from the subsidy scheme. It also replaced the bonus depreciation deduction system by a non-refundable fixed-rate investment tax credit of 25 percent of the purchase price (5 percent in the first year, then 2.5 percent a year for eight years), which directly reduced an investor’s tax liability instead of offsetting taxable income. An alternative scheme, the “social” *Scellier* law, similar to the earlier *Borloo* scheme, allowed households to add to the basic *Scellier* tax benefits

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<sup>74</sup>Successively: *Perissol*, *Besson*, *Scellier*, *Robien*, *Duflot*, and *Pinel*. Historian and sociologist Vergriete (2013) provides a long-run view of the political implications of these policies.

an option for a 30 percent deduction off rental income, under the condition that rents be capped c. 20 percent below *Scellier* rent ceilings. Due to budget constraints amid the euro area financial crisis, subsidies to newly produced rental housing investment under the *Scellier* and *social Scellier* schemes were first sharply reduced in 2011: the rate of the investment tax credit was lowered to 22 percent for units respecting strict environmental and energy efficiency, and to 11 percent for other units. The tax credit rate was then drastically cut down in 2012 to 13 percent for energy-efficient units, and to 6 percent for other units. Maximum purchase prices per square meter, which varied by zone, were also implemented starting in 2012.

- From January 2013 to September 2014, the *Duflot* scheme allowed for an 18 percent tax credit (2% annually over nine years), but implemented stringent rent conditions, and excluded B2 and C cities (unless they obtained an agreement from the prefect) from the benefit of the scheme.
- Finally, after September 2014, the *Pinel* scheme reshuffled eligible areas, upgrading towns with c. 10 million inhabitants into zones with higher rent and income ceilings, and allowed for a more flexible rental commitment duration (from 6 to 12 years) and the possibility to rent a unit to one’s relatives. More information on the *Pinel* scheme is provided in the main text. Figure B.2 provides information derived from individual tax data on the take-up of the *Pinel* scheme, distinguishing purchases by the year in which taxpayers first claim the rebate (completion year) and by the delay since the purchase of the unit.



**Figure B.2: Total value of Pinel take-up by completion year** The figure plots the annual value of Pinel investments by the year they were first reported by tax filers – corresponding to the year the unit was first rented, along with the delay between the reporting year and the year the purchase was effectively made. For example, Pinel investments first reported for tax filing year 2017, but for which the purchase occurred in 2015, represent c. EUR 1.08 billion, while Pinel investments corresponding to purchases made in 2017 and rented for the first time in the same year were c. EUR 6.6bn. Sources: *Annuaire statistique* database, the annual report of the French Treasury reporting all personal income tax-related information - see appendix A for a detailed description of the data.

**Public interest in landlord incentives** Figure B.3 plots the evolution of interest in each of the last four schemes in operation (Robien, Scellier, Duflot, and Pinel) since 2004. Interest for each of the schemes (as measured by GoogleTrends search volume, normalized to 100 at the maximum value) follows its implementation and timing of availability. Within the period in which schemes are available, interest gauged by Google searches spikes towards the end of each fiscal year, and close to planned eligibility deadlines, demonstrating the salience of tax motives, and consistent with empirical evidence on the actual timing of purchases which tend to bunch towards the end of fiscal years and immediately before the end of eligibility.

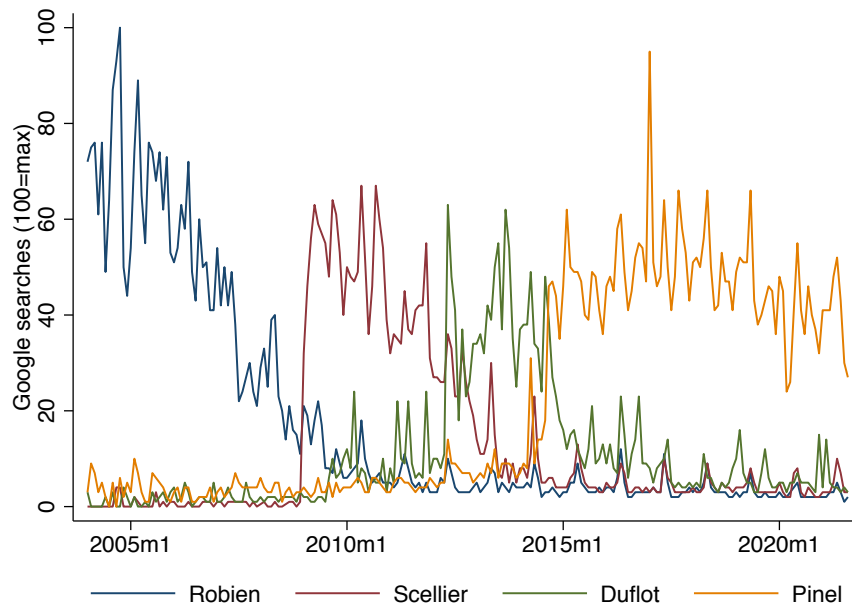
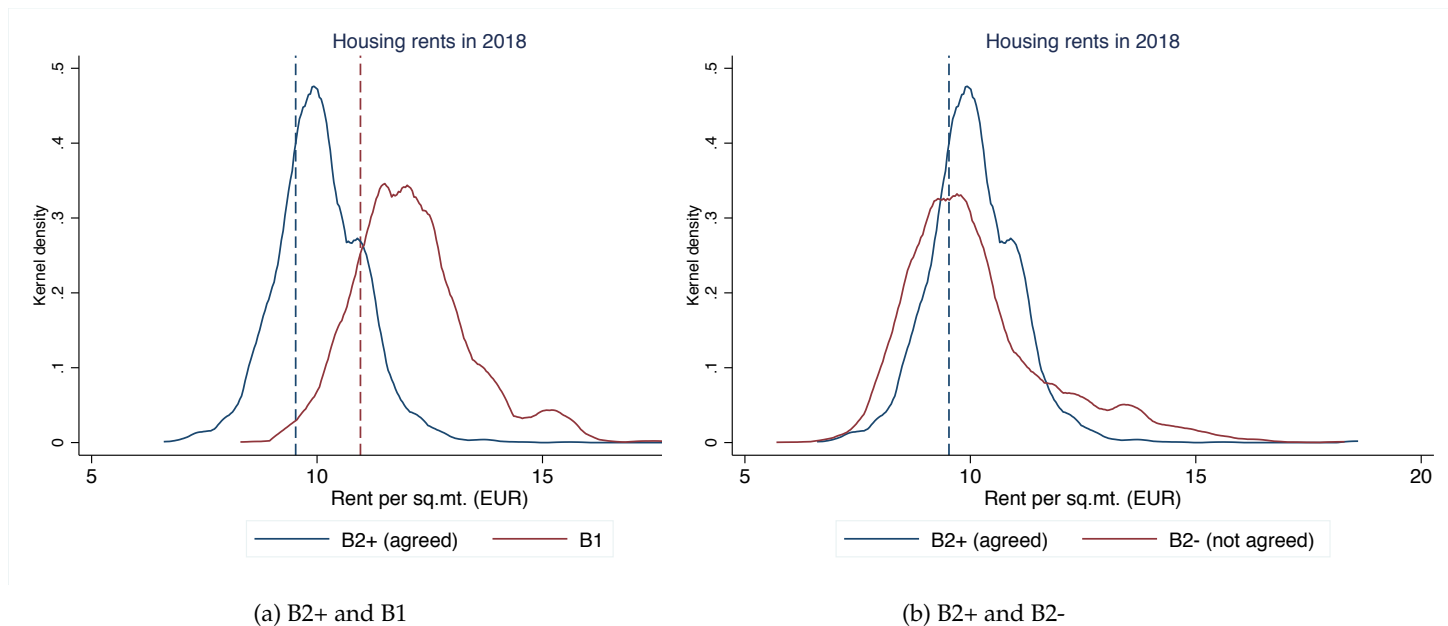


Figure B.3: **Interest in tax-favored schemes** The figure plots the Google Trends index of search intensity (normalized to a maximum value of 100) for the last four schemes in operation (Robien, Scellier, Duflot, and Pinel). The successive tax regimes are described in table 4.

**Net present value of the Pinel incentive** Rents under the Pinel scheme are theoretically subject to a ceiling amount, designed to target a 10 to 20 percent discount relative to average market rents in a zone. In 2017, it was set at  $C = \text{EUR } 16.83/\text{sq.m.}$  in area A bis;  $\text{EUR } 12.5/\text{sq.m.}$  in A;  $\text{EUR } 10.07/\text{sq.m.}$  in B1;  $\text{EUR } 8.75/\text{sq.m.}$  in B2. For example, a *Pinel*-eligible 50 sq.m. apartment in Marseille (zone A) would be subjected to a  $C \times M = 12.5 \times (0.7 + 19/50) = 13.5 \text{ EUR}/\text{sq.m.}$  ceiling, while a *Pinel*-eligible 30 sq.m. apartment in Toulouse (zone B1) would face a  $C \times M = 10.07 \times 1.2 = 12.1 \text{ EUR}/\text{sq.m.}$  ceiling. Resident household income was capped (for a couple with one child) at 72,737 € in area A bis; 66,699 € in A; 48,596 € in B1; and 43,737 € in B2. Area-specific income ceilings vary with family size, with larger households subject to looser maximum income guidelines. The income ceiling only applies in the first year of the lease. Nationwide, the median household income for a couple with one child was 43,880 € in 2017. Income ceilings cover close to 80 percent of the population and rarely bind.

The effective value of the *Pinel* subsidy to new investment varies across locations depending on the depth of

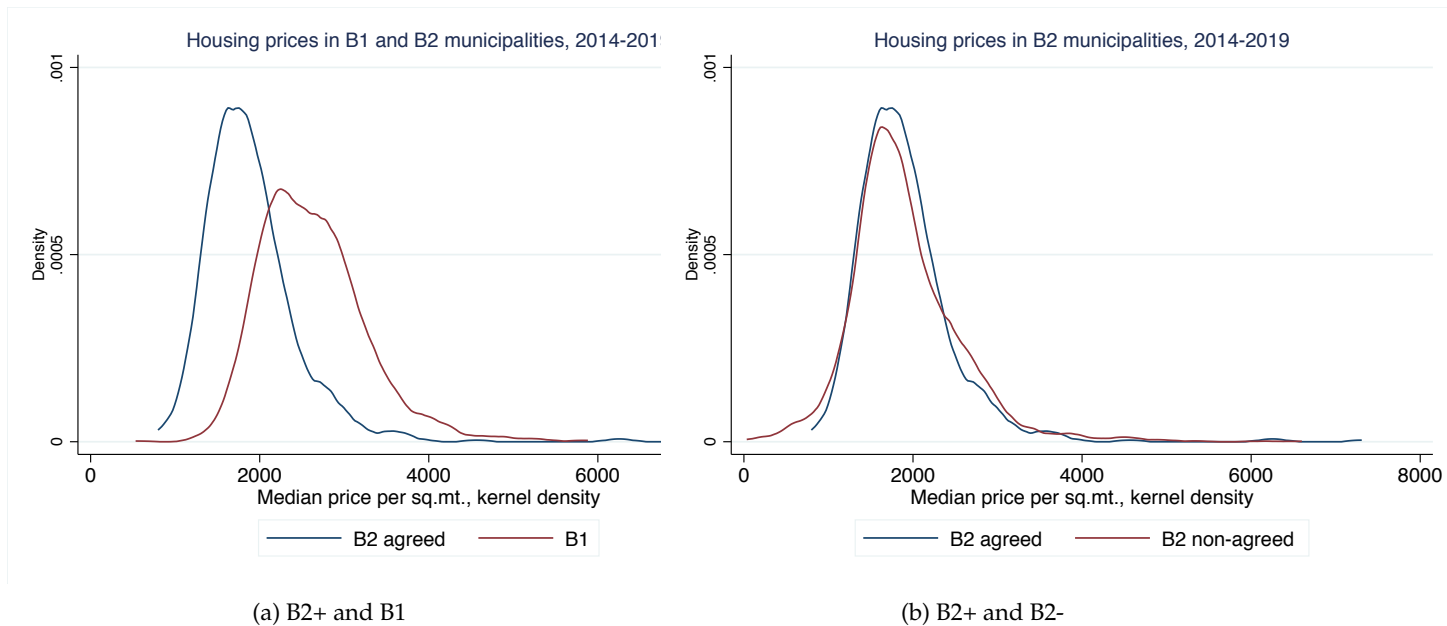
required rent discounts relative to market rents. Nonetheless, using data on effective market rents and ceilings applicable under the Pinel scheme, it is possible to estimate a range of possible values for the effective net present value of the incentive embedded in the Pinel scheme. I make data-driven but highly conservative assumptions for the discount rate, the extent to which affordability requirements fall below market rents, the duration of investors' commitment to affordability, and average market rent-to-price ratios. I take a baseline rent-price ratio of 5 percent, close to the median in the sample of B2+ municipalities. During the period of interest, interest rates on 10-year French government bonds (the relevant discount rate for government-guaranteed future transfers) were in the range of 1 to 2 percent; this value is similar to the interest rates on new mortgages documented by the Bank of France over the period. Among B2+ municipalities, the mean rent markdown implied by the *Pinel* regulation relative to baseline market rents was 6 percent, and the 90<sup>th</sup> percentile was 17 percent (see figure B.4), so that central estimates of the NPV of the incentive should cluster around values around 5 to 15 percent for the depth of the rent markdown. Lower baseline rent-price ratios, a 12-year commitment, or non-binding rent markdowns would deliver even larger values of the NPV of the incentive. Table 3 provides a sensitivity analysis for the value of the tax incentive net of affordability requirements. While preferred estimates are in the upper-right quadrant of the table, even for assumptions of very high discount rates and a severe bite of rent discounts, the net value of the incentive is still above or close to 8-9 percent of the initial purchase price of the asset.



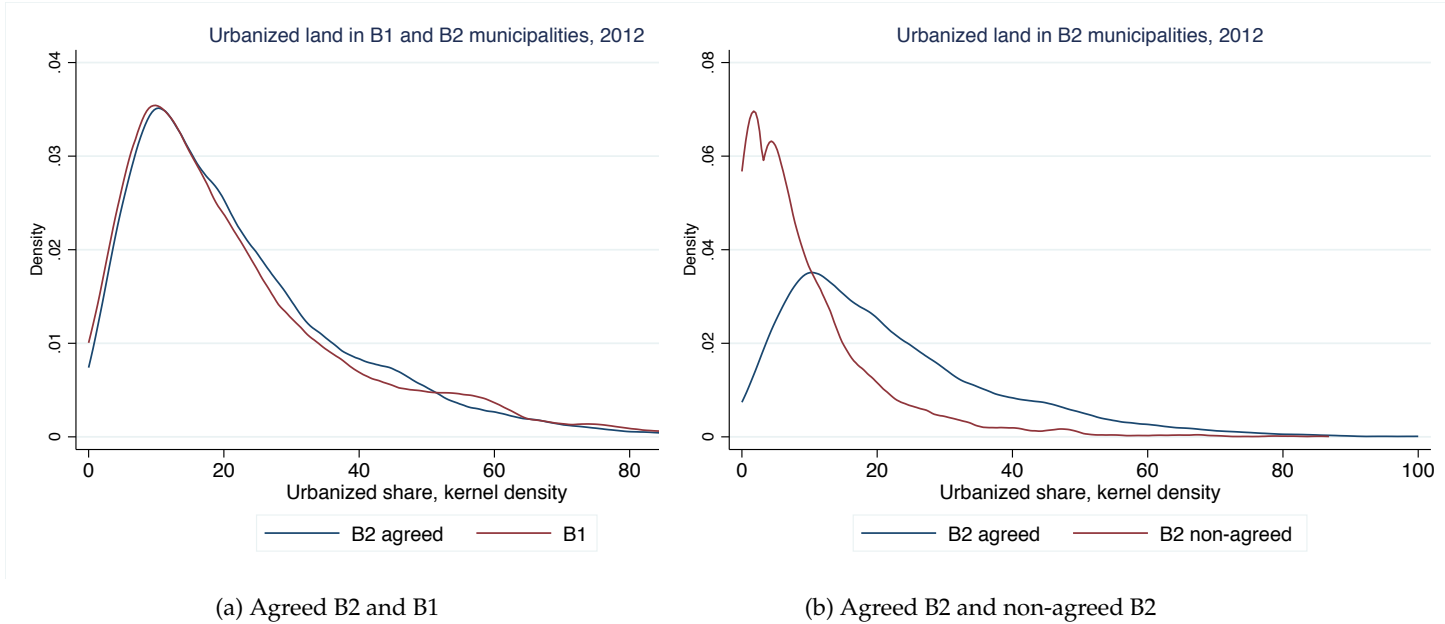
**Figure B.4: Rents by zoning area** The figure plots the rents per square meters for apartments in 2017-2018 for cities in the main treatment group – B2+ municipalities with an agreement (blue line), and their counterparts used as controls – always-treated B1 and never-treated B2- municipalities. Vertical dashed lines mark the rent ceiling (for a 50 sqm apartment) under the *Pinel* scheme for each area's corresponding color. Rents display substantial heterogeneity within each area, and substantial overlap across areas. Sources: French Housing Secretariat – see appendix A for a detailed description of the data.

## B.2 Legal ranking of municipalities by property market tension

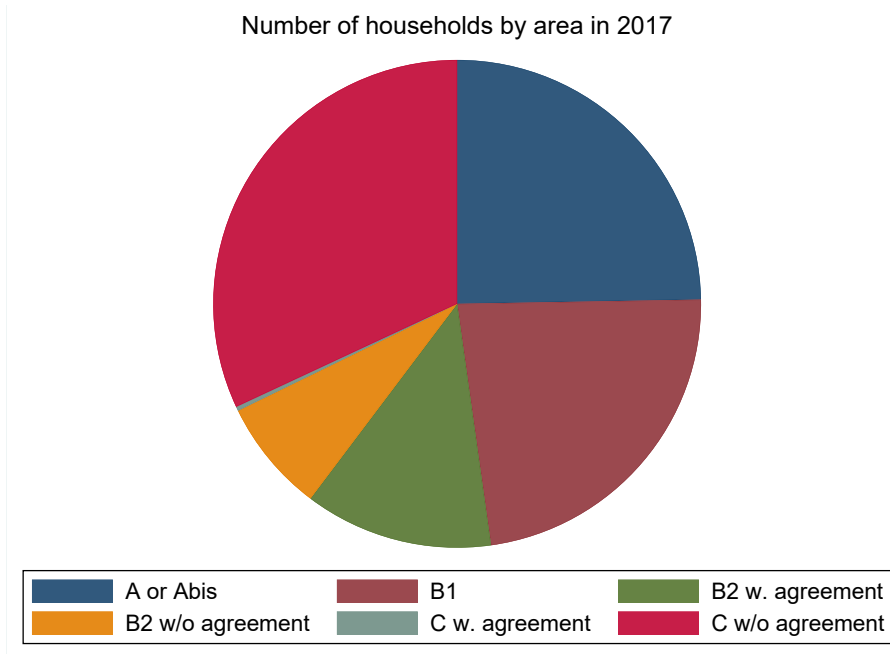
Higher housing price markets were broadly allocated to higher-ranked areas. One may think of the classification as a slightly more granular version of the United States' "Difficult Development Area" status, whereby areas with higher housing costs are ranked at a higher degree of priority for government support to low-income housing. Nonetheless, there was both substantial heterogeneity in housing demand and supply conditions *within* each one of the five areas, and significant overlap *across* them. As an example, figure B.5, panel (a) compares the (kernel density smoothed) distribution of municipality-level median home prices from 2014 to 2019 for municipalities in the B1 and B2+ areas; panel (b) of the same figure compares agreed municipalities in the B2 zone with B2 municipalities without an agreement (B2- towns). The distribution of median house prices by municipality closely track each other in agreed and non-agreed B2 municipalities, although there is substantial overlap between the agreed B2 and B1 municipalities. Figure B.6 does the same comparisons for the share of urbanized land. This time, the distribution of urbanized land cover by municipality more closely resemble each other in B2+ and B1 municipalities, although there is also substantial overlap between the B2+ and B2- cities.



**Figure B.5: Median house prices by zoning area** The figure plots the median home price for apartments and houses sold from 2014 to 2019 available in the DVF database for cities in the main treatment group – B2 municipalities with an agreement (blue line), and their counterparts used as controls – either B1 (panel (a)) or B2 (panel(b)) municipalities (red line). Home prices display substantial heterogeneity within each area, and substantial overlap across areas. Sources: *DVF* database, an exhaustive record of housing transactions in France made available by the French Finance Ministry. Sales restricted to houses and flats with an available transaction value and square footage - see appendix A for a detailed description of the data.



**Figure B.6: Share of urbanized land by zoning area** The figure plots the share of urbanized land available in the Corine Land Cover database for cities in the main treatment group – B2 municipalities with an agreement (blue line), and their counterparts used as controls – either B1 (panel (a)) or B2 (panel(b)) municipalities (red line). Urbanized shares display substantial heterogeneity within each area, and substantial overlap across areas. Sources: *Corine Land Cover 2012* database, a land cover inventory initiated in 1985, providing detailed information on the use of local land (land cover) broken down into 44 classes. Data are made available at the municipality level by the French Environment secretariat.



**Figure B.7: Population by Pinel zoning** This figures provides the breakdown of household numbers (as of 2017) by municipality type. Municipalities were mapped into one of five areas, determining both their eligibility to the *Pinel* scheme, and rent guidance and income restrictions. Areas are ranked in descending order of estimated property market "imbalances", with Abis being the most highly-demanded areas and C being the least demanded. B2 and C areas are further broken down by agreement status (B2+/C+ are B2 or C towns with a prefect's agreement that are eligible to the *Pinel* scheme).

### **B.3 The “garrison towns” special treatment**

The *Pinel* policy was discontinued for B2 municipalities with an agreement after January 1, 2018 for new projects (March 2019 for sales). However, the same legislation expanded the policy coverage to include all municipalities that were (or had been part at some point within the last 8 years) part of a program to revitalize former defence garrison towns which had lost military units due to a reorganization of French forces. These towns were part of contractual agreements between the Ministry of Defence and local authorities labelled CRSD or *Contrats de Revitalisation des Sites de Defense*. After January 2018, any town that had been recently covered by a CRSD plan was eligible to the *Pinel* scheme, independently of its classification. Therefore, CRSD-covered towns in C or B2 zones without an agreement prior to 2018 gained eligibility, while CRSD-covered B2+ towns kept their eligibility to the *Pinel* scheme in spite of the 2018 reform. For A-Abis-B1 towns, which all kept their eligibility to the scheme after 2018, the CRSD expansion did not affect the availability of *Pinel* scheme investments.

## C Stylized facts: additional evidence

### C.1 Rental housing and mobility

**Labor mobility and private rental housing are closely associated** I document that residential mobility and private rental housing are tightly linked, suggesting a spatial equilibrium framework should focus on the provision of rental, rather than owner-occupied, housing.<sup>75</sup> Using Census data aggregated at the municipality level, panel (a) of figure C.8 evidences that, in the (population-weighted) cross-section of France's 35,000 towns in 2018, a larger share of private rental properties in the local housing stock of a municipality is tightly correlated with higher housing turnover (as measured by the share of units which households moved in less than two years earlier). On average, a one standard deviation (10 percentage points) increase in the share of private rental dwellings in the local housing stock is associated with a 0.7 standard deviation (2.5 p.p.) higher share of households who moved in recently.<sup>76</sup> Using individual-level information on the distribution of tenure duration by status (renter versus owner-occupier) among mainland France's more than 29 million primary residences in 2019, panel (b) of figure C.8 shows that occupancy in the private rental sector is substantially shorter and left-skewed, relative to owner-occupied housing.<sup>77</sup>

While turnover at the dwelling level is relevant to housing market fluidity, some of the differential mobility of renters could be driven by a higher frequency of within-town moves. Using detailed town-to-town mobility data from the Census, panel (c) of figure C.8 examines longer-range geographic mobility, which is more relevant to the spatial reallocation of labor. It displays the close association between the degree of inwards migration in a town and local private rental dwellings availability.<sup>78</sup> In addition, not only do towns with a higher proportion of private rentals receive more movers, but inwards migration there comes from further away. Weighting bilateral inwards moves in a municipality by the distance between the previous residence and the current town, panel (d) of figure C.8 shows that towns with more private rental dwellings receive longer-distance inwards movers. Therefore, both the extensive (share of movers) and intensive (average distance among movers) margins of spatial mobility closely co-move with the availability of private rental housing in a location. These regularities motivate a model where "grounded" owner-occupiers are relatively immobile across space, but "footloose" renters move in response to - and partially arbitrage - utility differentials across

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<sup>75</sup>Spatial equilibrium models often abstract from differences in tenure status, but Blouri, Büchler, and Schöni (2021) and Favilukis and Van Nieuwerburgh (2017) both characterize residential tenure mode as an endogenous household decision. Fixed transition costs in and out of owner-occupied housing can be large, making renters more prone to move. Monetary fixed costs specific to homeowner mobility include real estate transfer taxes, realtor fees, or capital gains taxes after a nominal price appreciation (see e.g. Levy (2021) for quasi-experimental evidence of the impact of the latter in France). Other, non-pecuniary, differential fixed costs may include deeper local ties for homeowners or behavioral biases such as endowment effects. The higher mobility of renters documented in this section could thus be due either to a *causal* effect of tenure (e.g. through differential mobility costs), or *selection* patterns, as households more prone to move due to observable and unobservable characteristics tend to choose renting rather than owning (Oswald, 2019).

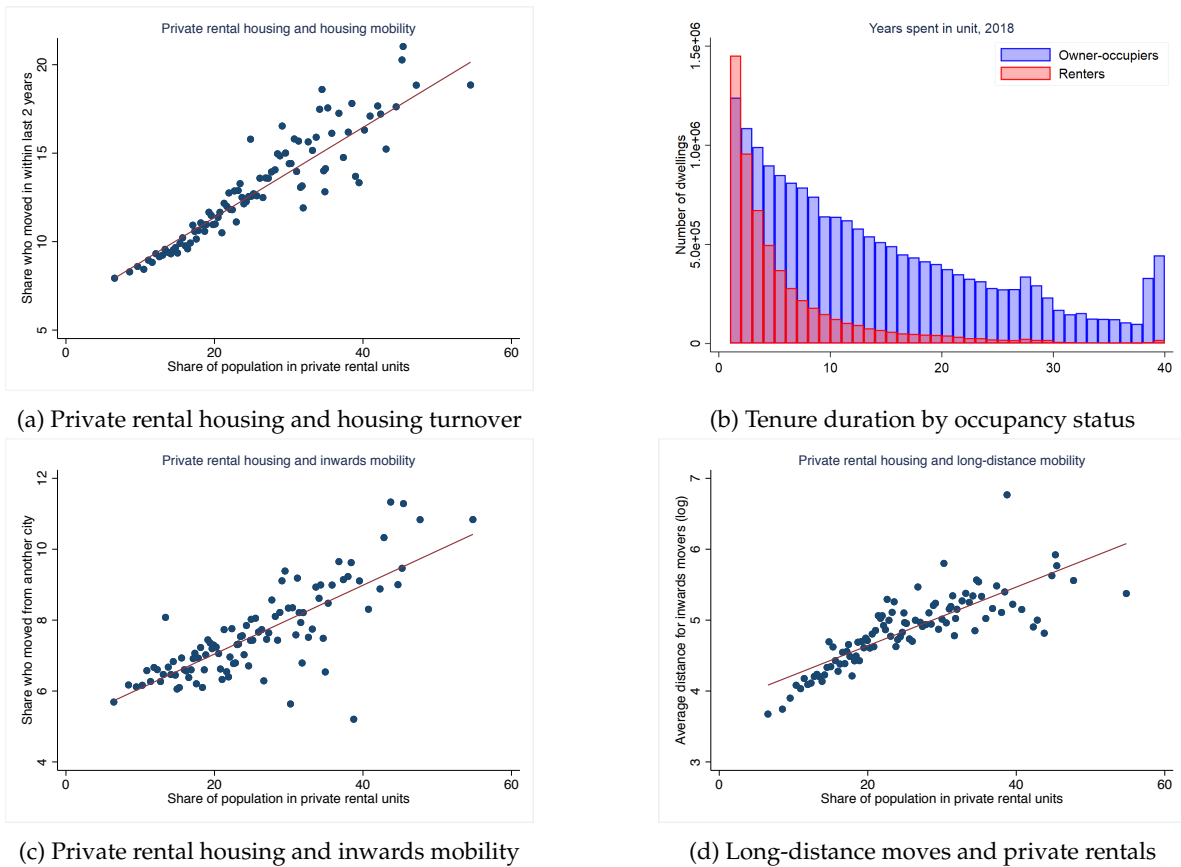
<sup>76</sup>On the other hand, the correlation of this measure of household mobility with the share of *public* rental housing is almost nil (panel (E.1a) of appendix figure E.1). Public housing (provided by the government at deep discounts against market rents) represents close to 40 percent of all rental units in France, or close to five million units.

<sup>77</sup>The median tenure for renters in the private sector is 2 years [interquartile range: (0,5)], while the median duration for owner-occupied units is 10 years [interquartile range: (4,20)], and 5 years [interquartile range: (2,12)] in *public* rental housing.

<sup>78</sup>Panels (b) and (c) of appendix figure E.1 show no such close linkage exists between cross-city moves and the share of *public* rentals in the local housing stock, as both tend to be *negatively* correlated. The same regularity holds for outwards migration, which is strongly and tightly positively correlated with the availability of private - but not public - rental housing.



locations, making the availability of *rental* housing a key determinant of overall mobility.<sup>79</sup>



**Figure C.8: Private rental housing and spatial mobility** Panel **a** plots the (population-weighted) relationship between the share of the population living in private rental units in a municipality, and the share of the population living in units which they moved in less than two years earlier, in 2018. Panel **b** plots the distribution of tenure duration (measured as the number of years since the household moved into the property) separately for renters in the private sector and for owner-occupiers. Same-year movers are not shown, and the maximum tenure duration is censored at 40 years for legibility purposes. Panel **c** plots the (population-weighted) relationship between the share of the population living in private rental units in a municipality, and the share of the population which moved into the municipality in the last year, in 2018. Panel **d** plots the (flow-size weighted) average distance of moves ( $\bar{d}_i = \sum_j \frac{F_{ij}}{\sum_j F_{ij}} d_{ij}$ ) by inwards movers from  $j$  to municipality  $i$  in the last year (in logs), against the share  $s_i$  of the population living in private rental units in  $i$ , in 2018. The distance of moves is computed as the haversine distance between the centroids of the municipality of origin and the destination municipality. Panels **a**, **c**, and **d** present binned scatter plots where the average value of the y-axis variable is plotted against each (population-weighted) centile of the x-axis variable, and data for these panels come from France’s national statistical institute *Base Logements*, a housing survey matched to the Census and combined with register data. Data for panel **b** are computed from register data (the *FIDELI* database) on the universe of c. 29 million dwellings used as primary residences in France.

The differences in spatial and residential mobility behavior between private sector renters and owner-occupiers correlate with distinct observable characteristics of agents choosing each of the two tenure modes. Table C.1 provides demographic information on the reference individual for households in three subgroups (owner-occupiers, renters in the private sector, and renters in the public sector), obtained from exhaustive Census data. Private sector renters live in smaller households, are younger, more likely to be full-time employed and less likely to be retired, and are overwhelmingly more likely than either owners or public sector renters to have moved in recent years and to have lived in their home for a short time period.

<sup>79</sup>The close linkage between rental housing and mobility is not specific to the French context. In the United States, according to the Census Bureau for the year 2017, renters moved at an annualized frequency of 21.7 percent, against only 5.5 percent for owner-occupiers.

Table C.1: Characteristics of household by tenure mode

	Owner-occupiers	Private sector renters	Public sector renters
<b>Household size</b>	2.30 (1.21)	1.89 (1.17)	2.30 (1.43)
<b>Age of reference individual</b>	58.7 (16.5)	43.6 (18.2)	51.3 (16.8)
<b>Full-time employed</b>	0.56 (0.50)	0.61 (0.49)	0.52 (0.50)
<b>Out of labor force or retired</b>	0.40 (0.49)	0.20 (0.40)	0.31 (0.46)
<b>Unemployed</b>	0.04 (0.19)	0.12 (0.32)	0.16 (0.37)
<b>In current home for 5+ years</b>	0.82 (0.38)	0.34 (0.48)	0.64 (0.48)
<b>Moved since last year</b>	0.05 (0.22)	0.25 (0.43)	0.1 (0.3)
<b>Weighted N</b>	17,107,944	7,585,030	4,390,937

Statistics are computed from the 2018 full Census microdata. The sample is restricted to the reference individual for all households in each of the three tenure modes, excluding respondents in non-traditional or institutional housing. Estimates use sampling weights provided by the national statistical institute INSEE, and provide the mean and (in brackets) standard deviation for each variable.

All of these observable characteristics are also associated with a more substantial propensity to move across municipalities. Nonetheless, when estimating a linear probability model for having moved from a different municipality in the last year on a variety of individual characteristics, as presented in table C.2, the coefficient on renting remains high, even after controlling for a number of individual covariates. The partition of the population of households alongside characteristics associated with spatial mobility indicates a clear segmentation between the population of mobile private sector renters and mostly immobile owner-occupiers.

Spatial variation in the propensity to move is closely associated with the presence of private rental housing in a given location, as documented above for the cross-section of French municipalities. This fact also holds over time within cities. Using panel data from 2009 to 2018 at the municipality-year level obtained from the French Housing survey matched to Census data, I show in table C.3 that even conditioning on municipality fixed-effects, a higher share of private rental housing is tightly associated with increases in measures of spatial and housing mobility. This pattern is not claiming a direct causal link from the availability of rental housing to the evolution of household mobility, but the existence of a residual correlation within municipalities over time provides suggestive evidence that the close connection between the two phenomena captures more than time-invariant structural characteristics of a locality.

Table C.2: Probability of having moved across cities

	(No controls)	(With controls)
Owner-occupier		(Reference category)
Private sector renter	0.14**** (0.00)	0.07**** (0.00)
Public sector renter	0.01**** (0.00)	-0.01**** (0.00)
R-Square	0.05	0.09
Observations (unw.)	8762383	8762383

Standard errors in parentheses. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$

Regressions coefficients are obtained from a linear probability model using the 2018 full Census microdata. Controls in the second column include household size, an indicator for being full-time employed, an indicator for being out of the labor force or retired, and a quadratic in the age of the reference individual. The sample is restricted to the reference individual for all households in each of the three tenure modes, excluding respondents in non-traditional or institutional housing, corresponding to a weighted number of observations of  $N^w = 29,083,911$ . Estimates use sampling weights provided by the national statistical institute INSEE, and robust standard errors to heteroskedasticity.

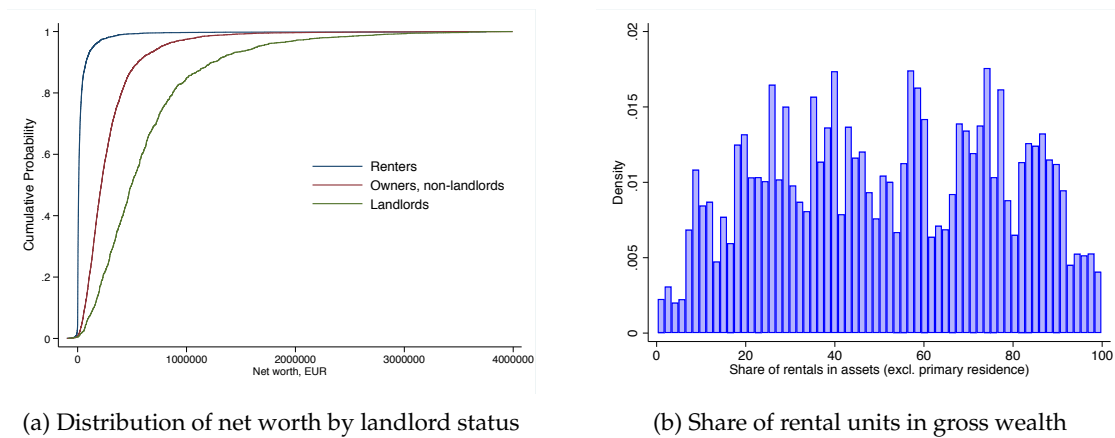
Table C.3: Recent moves and private rental share

	Pooled	Year FE	City FE	2-way FE
Share of private rentals	0.363**** (0.0327)	0.363**** (0.0328)	0.150**** (0.00997)	0.162**** (0.0107)
R-Square	0.572	0.573	0.936	0.937
Observations	359390	359390	359388	359388
Clusters	36682	36682	36680	36680
Municipality FE	No	No	Yes	Yes
Year FE	No	Yes	No	Yes

Standard errors in parentheses. \*  $p < 0.10$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ , \*\*\*\*  $p < 0.001$

Regressions coefficients are shown for the association between the share of private rental housing in a municipality and the share of households who moved in less than two years ago. They are obtained from a panel regression with city and year fixed-effects, across France's municipalities, using data from the French Housing survey matched to the Census for 2009-2018. Estimates use robust standard errors clustered at the municipality level, and are weighted by city population.

## C.2 Housing investor characteristics



**Figure C.9: Landlords wealth and limited diversification** Panel **a** plots the cumulative distribution function of net worth for households in three groups: renters, owner-occupiers who do not own any rental dwelling, and landlords defined as households who own at least one unit that they rent out for the entire year. Panel **b** plots the distribution of landlords according to the portfolio share of rental units in their wealth, excluding the value of their primary residence, using matched household-assets data from the wealth survey. The median portfolio share of rental units in gross wealth is 52 percent. Data are computed at the household level, from the *Enquete Patrimoine* 2017-2018, a wealth and income survey matched to summary information from household tax returns which oversamples high wealth and high income individuals.

To motivate the segmentation between the saving behavior of landlords and the rest of the population, I show using microdata from a representative survey of households assets that landlords tend to differ substantially from both other owner-occupiers, and from renters, along numerous characteristics. Using wealth survey data from 2018, table C.4 provides some demographic information on the reference individual for households in three subgroups (landlords, non-landlord owner-occupiers, and renters), obtained from the *Enquete Patrimoine*. Landlords have higher gross wealth, receive higher disposable income, are older and more frequently married, and tend to live in larger households than either renters or non landlord owner-occupiers. Both types of owners have on average lived for much longer than renters in their current residence location. Figure C.10 shows that the income distribution of landlords tend to first-order stochastically dominate the income distribution of both non-landlord owners and of renters.

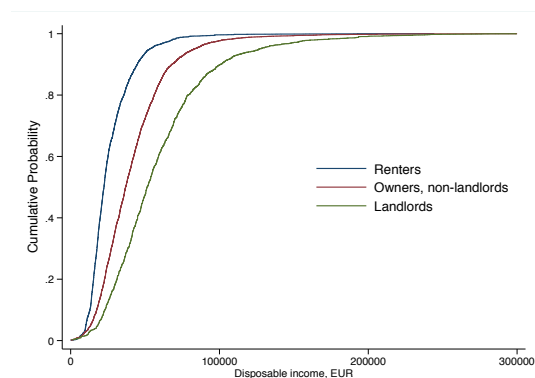


Figure C.10: Distribution of income by landlord status The figure documents the higher disposable income of landlords relative to other owner-occupiers who do not own rental housing and to renters. Data are computed at the household level, from the *Enquete Patrimoine* 2017-2018, a wealth and income survey matched to summary information from household tax returns which oversamples high wealth and high income individuals.

Table C.4: Characteristics of households by landlord status

	Landlords	Non-LL owners	Non-LL renters
<b>Household size</b>	2.52 (1.26)	2.38 (1.24)	2.05 (1.33)
<b>Age of reference individual</b>	55.8 (14.6)	57.6 (16.4)	48.8 (18.1)
<b>Gross wealth</b>	821,517 (1,605,048)	352,137 (897,508)	41,878 (172,445)
<b>Disposable income</b>	60,705 (53,649)	41,418 (30,443)	26,252 (15,477)
<b>Married</b>	0.74 (0.44)	0.67 (0.47)	0.34 (0.47)
<b>Owns more than one dwelling</b>	1 (0)	0.11 (0.31)	0.05 (0.23)
<b>Years in current dwelling (med.)</b>	15 (14.9)	17 (15.8)	6 (11.0)
<b>Weighted N</b>	2,958,885	13,988,108	11,209,854

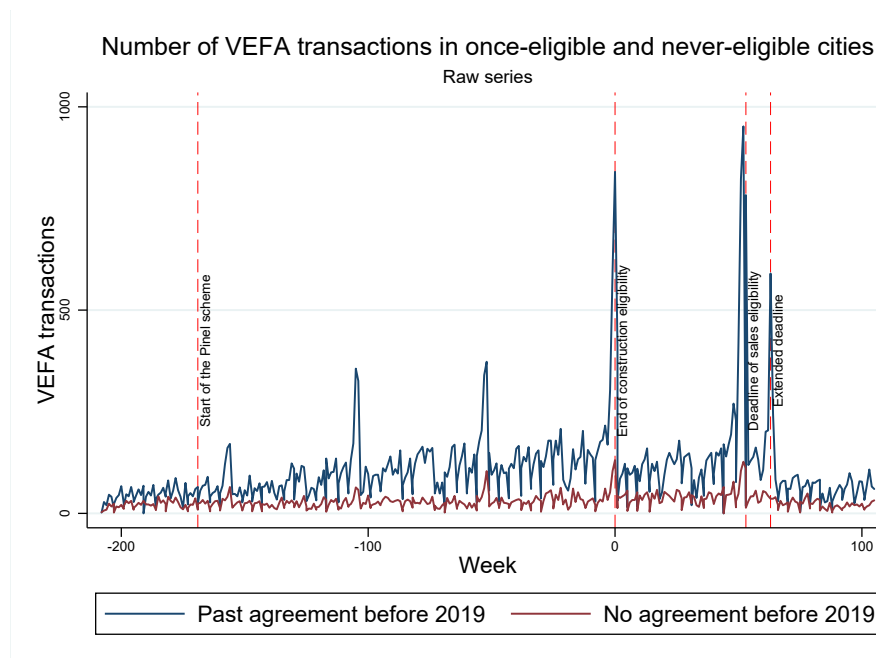
Statistics are computed from the 2018 *Enquete patrimoine* wealth survey. The sample is restricted to the reference individual for all households in each of the three subgroups, excluding respondents in non-traditional or institutional housing. Estimates use sampling weights provided by the wealth survey, and provide the mean and (in brackets) standard deviation for each variable.

## D Additional results

### D.1 Retiming and bunching evidence

I demonstrate graphically additional margins of adjustment to conditional housing subsidies. First, a high-frequency analysis of the timing of both developer requests of building licenses, and signing dates of actual transactions, demonstrates the salience and key role of tax motives in triggering housing supply decisions. Second, households bunch the purchase price of their new housing acquisitions below a kink in the subsidy schedule, providing evidence of sensitivity to changes in marginal subsidy rates.

**End-of-year abnormal investment behavior: sales** Abnormal end-of-year bunching in transactions is observed in the last month of each fiscal year for “off-plan” sales (denoted as *VEFA*, which correspond to pre-sales of units still under construction or for which construction has not started yet), even in the absence of any change in eligibility or tax treatment, as shown in figure D.11.



**Figure D.11: Total number of “off-plan” sales in B2+ and B2- municipalities** The figure plots the weekly number of sales identified with an “off-plan” dummy in France from 2014 to June 2020, respectively in never-eligible B2- (red line) and once-eligible B2+ (blue line) municipalities. Sources: *DVF* database, an exhaustive record of housing transactions in France made available by the French Finance Ministry. Sales restricted to houses and flats with an available transaction value - see appendix A for a detailed description of the data.

This provides evidence of the salience of the tax subsidy, but also potentially of a rational time-shifting of individual investment decisions in order to minimize the time elapsed between the investment in residential housing, and the claiming of the tax credit rebate during the next fiscal year. This behavior is particularly pronounced in anticipation of changes to the value of the subsidy in B2+ locations, and represents the household

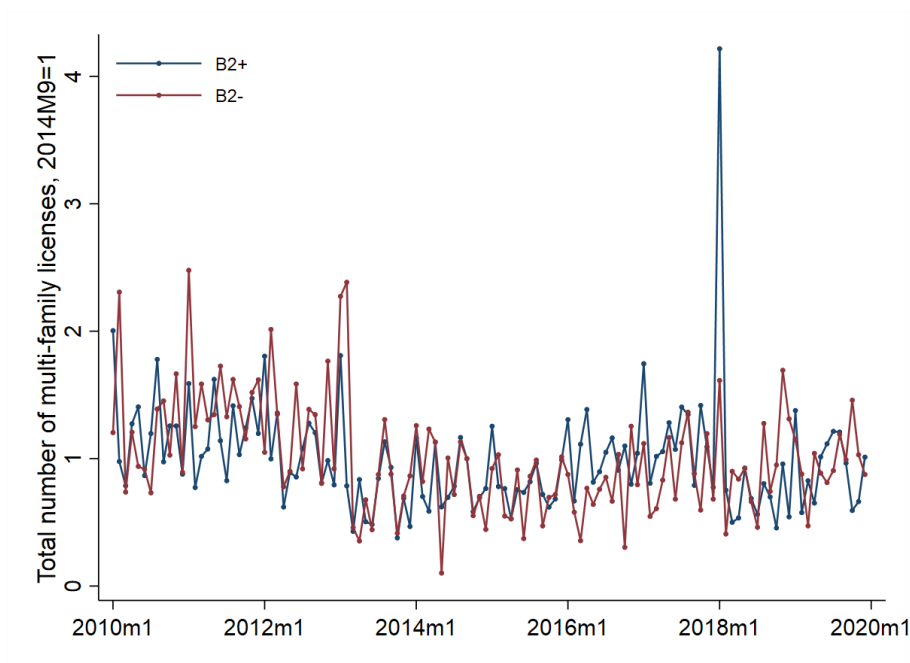


Figure D.12: **Total number of building license requests in B2+ and B2- municipalities** The figure plots the monthly number of building license requests for multi-family units, respectively in never-eligible B2- (red line) and once-eligible B2+ (blue line) municipalities. Sources: *Sit@del* database, an exhaustive record of building license requests - see appendix A for a detailed description of the data.

analog of the end-of-year bunching of tax-minimizing investments documented by Xu and Zwick ([forthcoming](#)) for the case of US firms subject to the corporate income tax. Sales in eligible B2+ cities exhibit sharp monthly spikes at the initial planned eligibility limit (December 2017), and at the interim (December 2018) and final (March 2019) eligibility deadlines, providing further credible evidence of the causal impact of the subsidy on new home sales.<sup>80</sup> These spikes are statistically and economically significant: in the last week of 2018, new home sales in the 900 eligible B2 municipalities reached more than EUR 200 million, roughly twenty times the average weekly new home sales in both eligible and ineligible cities after the policy ended.

**End-of-eligibility re-timing: licenses** For housing building permits granted to developers, the reform of the Pinel scheme (which was passed in Parliament in November 2017) allowed for housing sales in B2+ cities to remain eligible only if a building license for the unit had been *requested* before December 31, 2017. In response to this incentive, the number of building licenses requested by developers in eligible B2+ cities spiked in December 2017. I evidence in figure D.12 the presence of this spike in developer requests of building licenses immediately before the planned end of eligibility, using detailed information from the *Sit@del2* (building permits requests) database.

<sup>80</sup>Consistent with the findings of this paper, Berger, Turner, and Zwick (2020) and Best and Kleven (2018), studying temporary homeownership subsidies in the United States and transaction taxes exemptions in the United Kingdom, respectively, also find significant evidence of time-bunching right before eligibility to each of the policies ended.

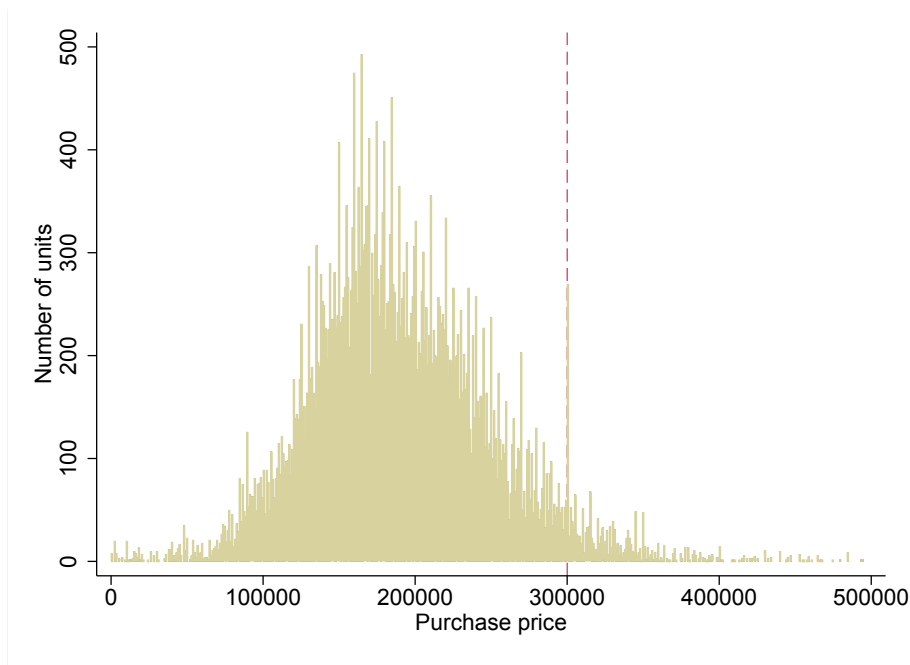


Figure D.13: **Purchase price of units under the Pinel scheme** The figure plots the distribution of purchase prices for units purchased under the Pinel scheme in the year 2016. Sources: *POTE* database, an exhaustive record of individual tax returns - see appendix A for a detailed description of the data.

**Kinks in the subsidy schedule: purchase price** Finally, exploiting exhaustive individual tax returns information, I evidence the presence of bunching at the kink in the subsidy schedule arising from the phase-out of the subsidy for units above 300,000 €. Figure D.13 plots the distribution of acquisition prices for all units under the *Pinel* tax schedule and evidences a clear spike exactly at 300,000 €, the value at which the proportional tax credit is maximized and above which the marginal subsidy is reduced to 0.

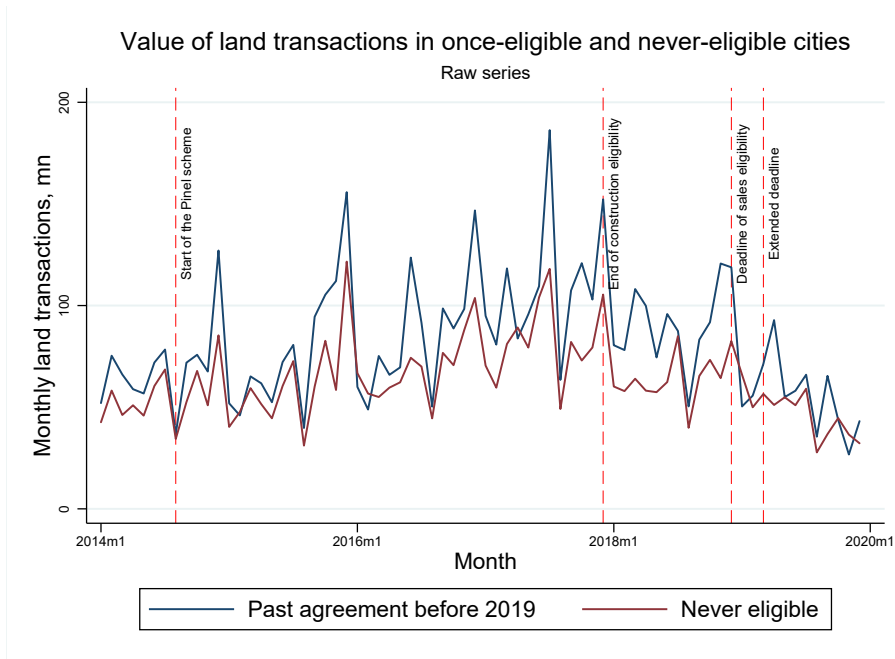
## D.2 Additional local effects of eligibility

**Land sales** A common concern surrounding housing subsidies is the capitalization of incentives in inelastically supplied local input prices, most notably land (Carozzi, Hilber, and Yu, 2019; Bono and Trannoy, 2019). Such input price rises push up construction costs as a consequence of housing subsidies. Using the DV3F database, I restrict the sample to sales of constructible land (so-called "*terrains a batir*").

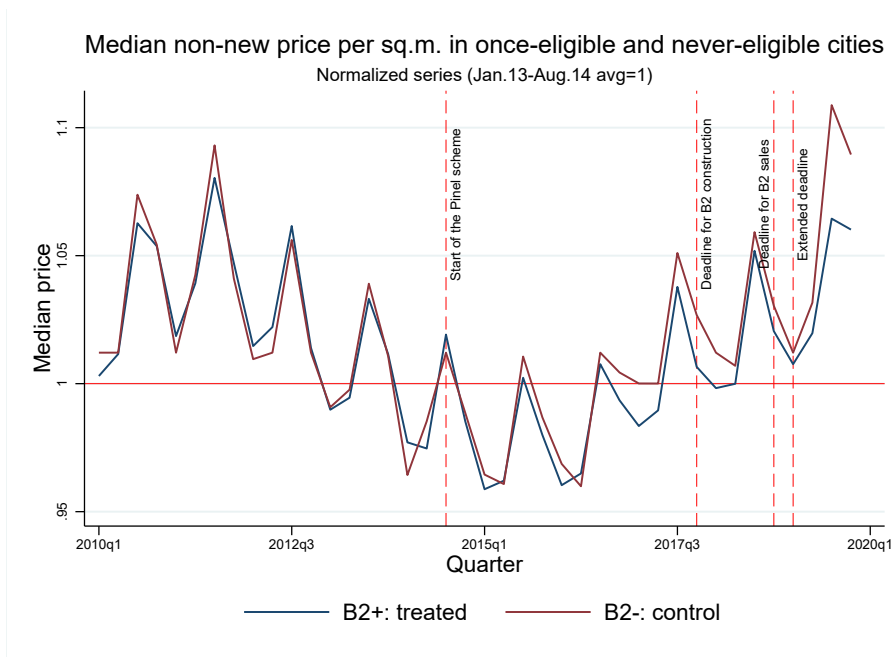
Figure D.14 plots the total monthly value of land sales in eligible B2 municipalities and their counterparts throughout the policy period. After following parallel trends to ineligible municipalities for much of 2014, the value of land sales in B2 towns with an agreement rose sharply throughout the policy eligibility period, especially in 2017, before the spike in licenses granted. After the end of the policy, the value of land sales in formerly subsidized B2 towns falls back to a trend and level comparable to ineligible cities.

**Prices of existing units** As stated by Kotlikoff (1983), a specific feature of investment tax incentives is their effect on the price of "old" capital: "*since equally productive units of new and old capital must sell for the same*

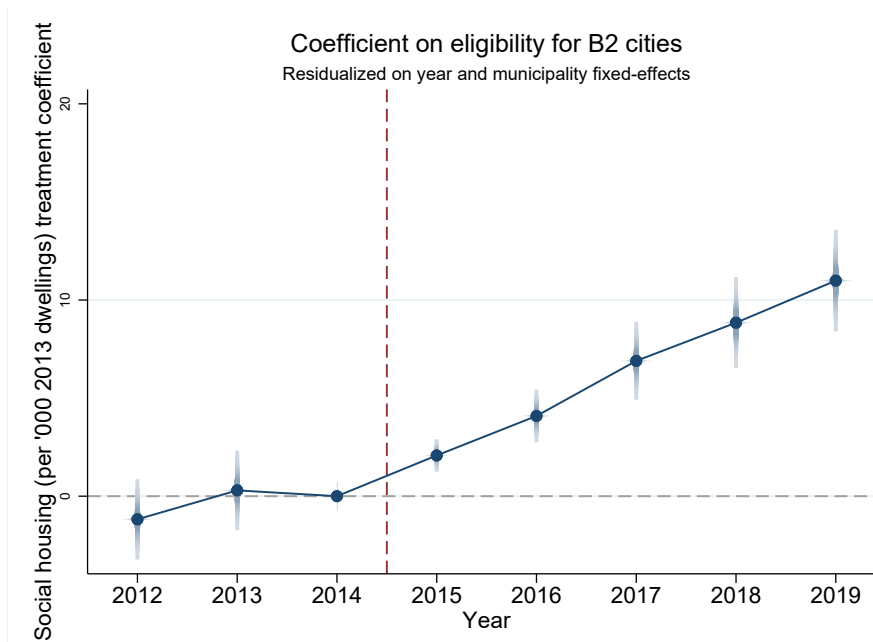




**Figure D.14: Total value of constructible land sales in B2 municipalities** The figure plots the monthly value of constructible land sales (million euros) in France from 2014 to June 2020, respectively in never-eligible (red line) and once-eligible (blue line) B2 municipalities. Sources: *DVF* database, an exhaustive record of housing transactions in France made available by the French Finance Ministry. Sales restricted to houses and flats with an available transaction value - see appendix A for a detailed description of the data.



**Figure D.15: Median price of existing units in B2 municipalities** The figure plots the median price per square meter in transactions for *existing* units, respectively in never-eligible (red line) and once-eligible (blue line) B2 municipalities. Sources: *DVF* database, an exhaustive record of housing transactions in France made available by the French Finance Ministry. Sales restricted to houses and flats with an available transaction value - see appendix A for a detailed description of the data.

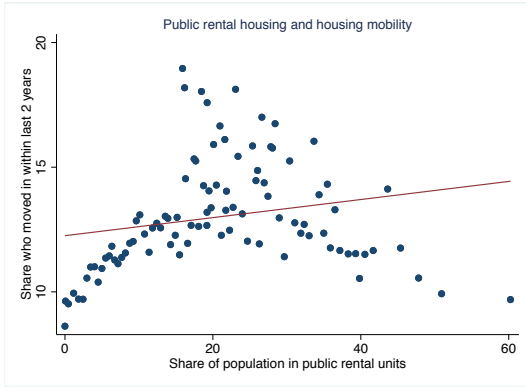


**Figure D.16: Effect on social housing units** The figure plots the estimated effect of eligibility to the *Pinel* scheme on social housing units. It plots, for B2+ (treated) relative B2- (control) towns, the coefficient on eligibility in equation 8 for the number of housing units per thousand of 2013 dwellings in the town. Standard errors are clustered at the municipality level and 95% confidence intervals are shown. Sources: *RPLS* database, an exhaustive record of social housing units in France made available by the French Housing Ministry – see appendix A for a detailed description of the data.

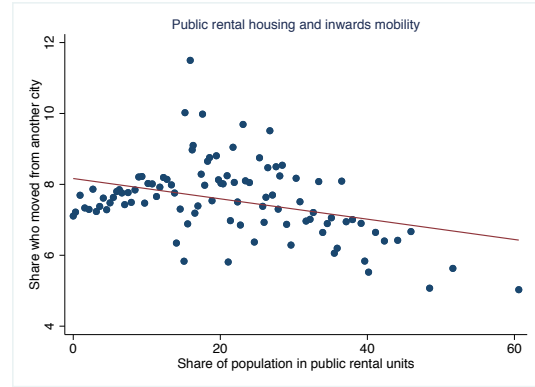
*price, tax provisions favoring new capital imply a lower price for existing capital*". While the only existing test of this hypothesis, to my knowledge, focuses on the price for used capital goods purchased by corporations (Edgerton, 2011), in my setting, observing local transactions for existing units in treated and untreated locations allows me to test for the presence of price effects for existing dwellings, i.e. location-specific used residential capital. Figure D.15 displays the normalized median price per square meter of existing units in B2+ and B2- municipalities over time, and provides suggestive evidence that treated locations experience a relative decline in the price of existing homes, as competing substitutable capital reduces their expected rental value.

**Social housing units** Using administrative data on the universe of social housing units provided by the French government, I also quantify a positive induced effect on social housing. Figure D.16 displays the results of the event-study regression coefficient on eligibility around entry. Part of the mechanism for the observed increase in social housing units stems from economies of scale, through which some of the individual units in a multi-family building targeted by developers for buy-to-let individual investors are purchased by local governments and public-private partnerships (*bailleurs sociaux*) to provide public housing in mixed-income communities.

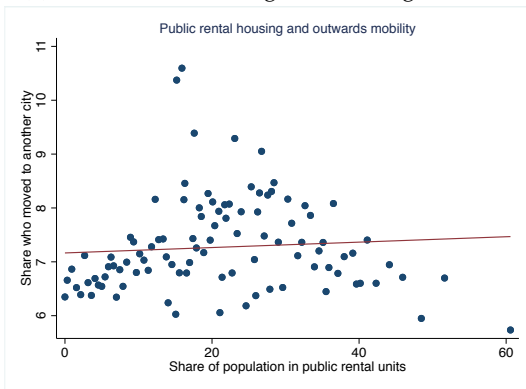
## E Additional appendix figures



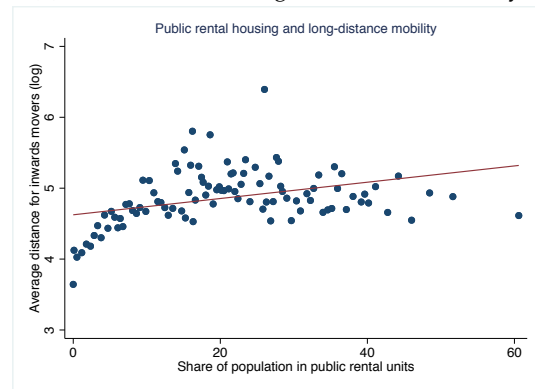
(a) Social rental housing and housing turnover



(b) Social rental housing and inwards mobility

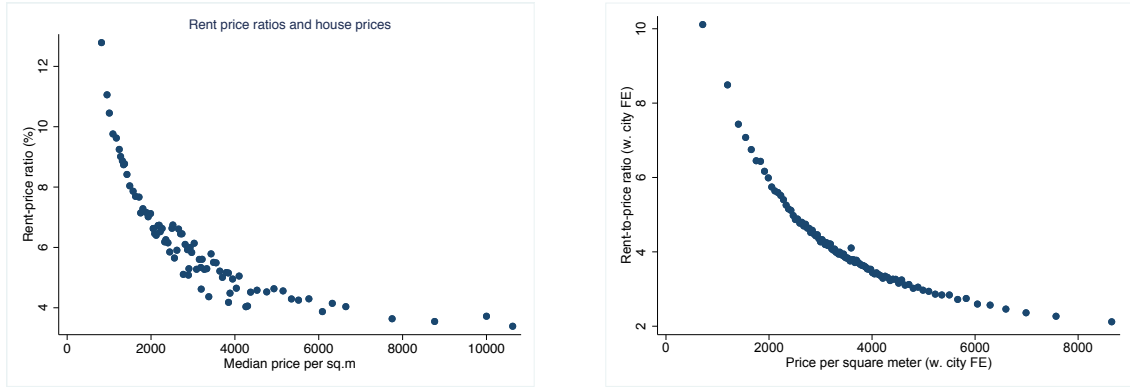


(c) Social rental housing and outwards mobility



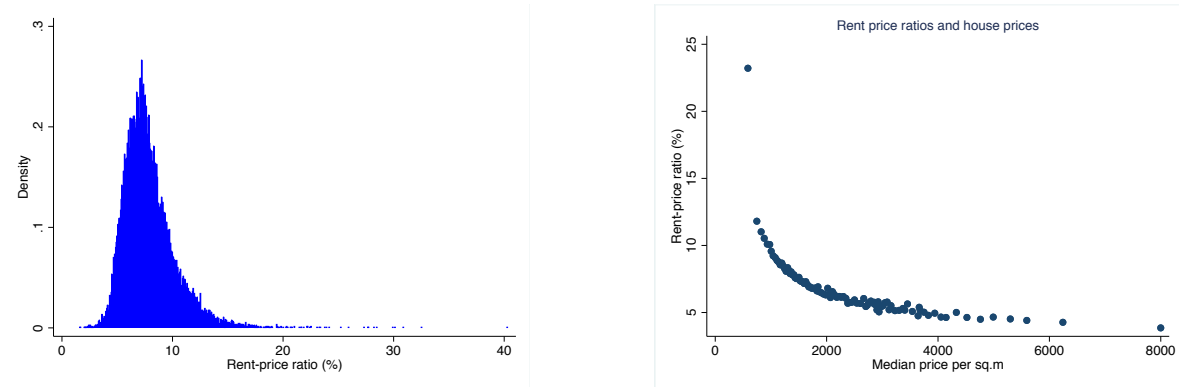
(d) Long-distance moves and social rentals

**Figure E.1: Public rental housing and spatial mobility** Panel **a** plots the (population-weighted) relationship between the share of the population living in public rental units in a municipality, and the share of the population living in units which they moved in less than two years earlier, in 2018. Panel **b** (resp **c**) plots the (population-weighted) relationship between the share of the population living in public rental units in a municipality, and the share of the population which moved into (resp. out of) the municipality in the last year, in 2018. Panel **E.1d** plots the (flow-size weighted) average distance of moves ( $\bar{d}_i = \sum_j \frac{F_{ij}}{\sum_j F_{ij}} d_{ij}$ ) by inwards movers from  $j$  to municipality  $i$  in the last year (in logs), against the share  $s_i$  of the population living in public rental units in  $i$ , in 2018. The distance of moves is computed as the haversine distance between the centroids of the municipality of origin and the destination municipality. All panels present binned scatter plots where the average value of the y-axis variable is plotted against each (population-weighted) centile of the x-axis variable, and data for these panels come from France's national statistical institute *Base Logements*, a housing survey matched to the Census and combined with register data.



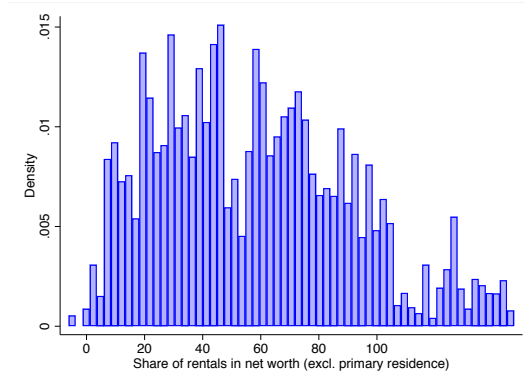
(a) Rent-price ratios and prices: across municipalities      (b) Rent-price ratios and prices: within municipalities

**Figure E.2: Rent-to-price ratios against prices** Panel **a** plots a binned scatter plot of rent-to-price ratios against median purchase price per square meter across municipalities, where y-axis values are averaged by centile bins of the x-axis variable. The rent-price ratios are computed using a cross-section of rents per square meter for multi-family units in 2018 made available by the French Housing secretariat at the municipality level, and median purchase prices for multi-family units from 2017 to 2019 using exhaustive transaction deeds data from the *DV3F* database. They only include municipalities with more than 11 transactions for multi-family units over the period. Panel **b** plots a binned scatter plot of rent-to-price ratios against purchase price per square meter across rental units purchased under the *Pinel* subsidy scheme, for municipalities with more than one unit, where both variables are demeaned by municipality, and y-axis values are averaged by centile bins of the x-axis variable. The rent-price ratios are computed using individual data on all c. 65,000 purchases made in 2016 under the *Pinel* scheme, exploiting buyers' commitment forms (2044-EB schedules attached to individual tax returns) which record the initial value of the rent under the affordable lease, as well as the purchase price and floor area of the unit to compute the taxpayers' tax reduction.

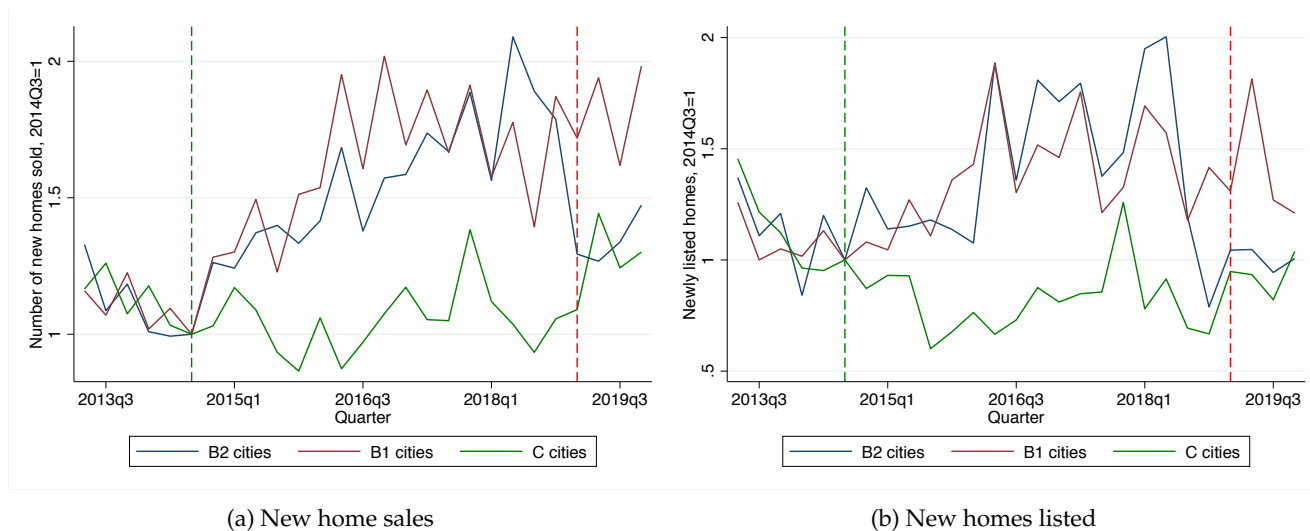


(a) Rent-price ratios: all municipalities      (b) Rent-price ratios and prices: all municipalities

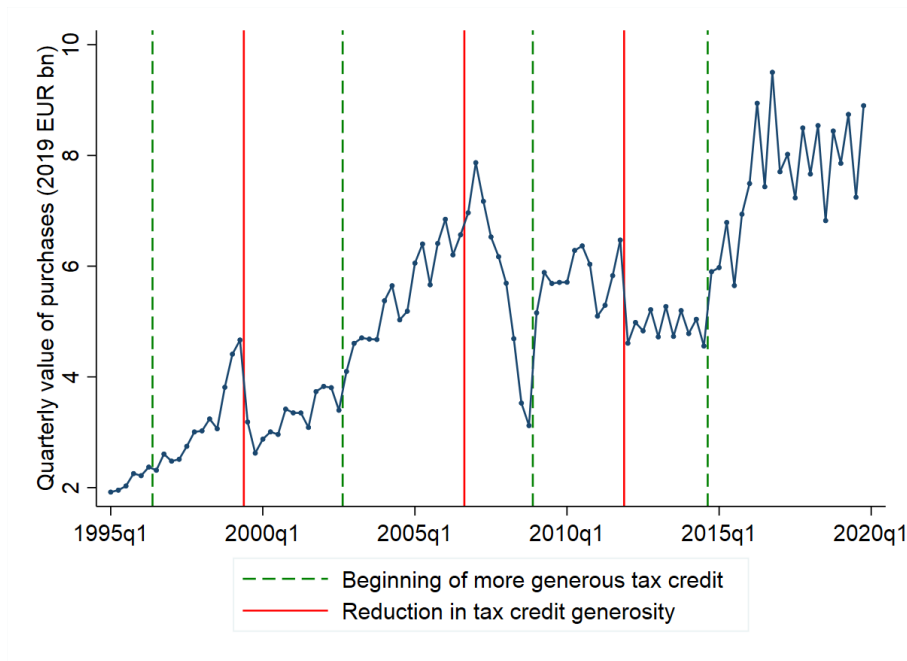
**Figure E.3: Rent-to-price ratios across all municipalities** Figure **E.3** documents the cross-city dispersion in rent-to-price ratios. It is constructed similarly to figure **2** but includes all municipalities with available data and does not restrict to municipalities with more than 11 transactions. Panels **E.3a** plots the distribution of rent-to-price ratios across municipalities. Panel **E.3b** plots a binned scatter plot of rent-to-price ratios against median purchase price per square meter across municipalities. The figures are computed using a cross-section of rents per square meter for multi-family units in 2018 made available by the French government at the municipality level, and median purchase prices for units from 2017 to 2019 using exhaustive transaction deeds data from the *DV3F* database.



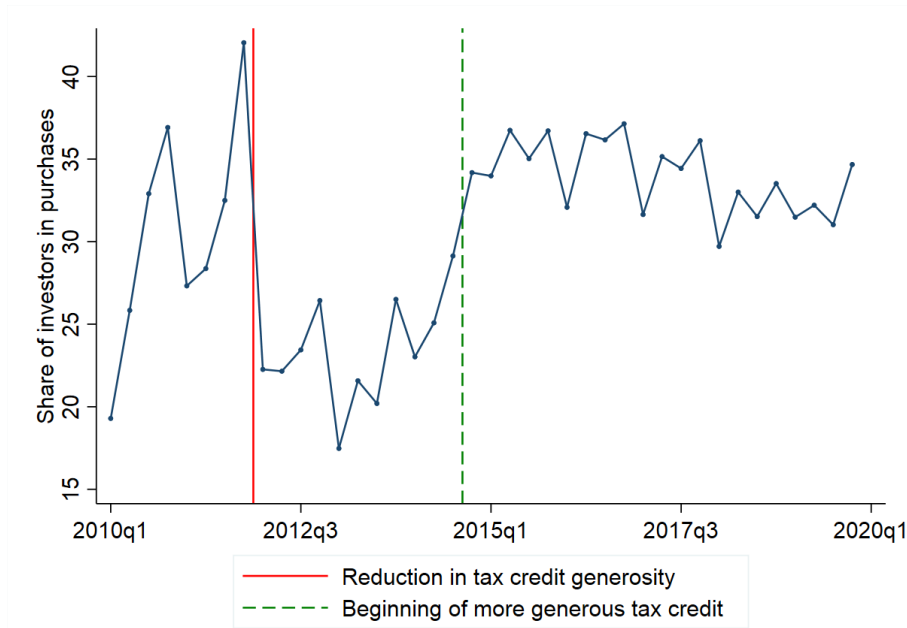
**Figure E.4: Portfolio share of rental units in net worth (excl. primary residence)** Figure E.4 documents the undiversified nature of housing investment in rental units. It plots the distribution of landlords according to the portfolio share of rental units in their net wealth, excluding the value of their primary residence, using matched household-assets data from the wealth survey. The median portfolio share of rental units in gross wealth is 52 percent. Data are computed at the household level, from the *Enquete Patrimoine* 2017-2018, a wealth and income survey matched to summary information from household tax returns which oversamples high wealth and high income individuals.



**Figure E.5: New home sales and new listings** Panel (a) plots the quarterly number of sales of newly built homes (normalized to 1 in 2014Q3) in B1 (blue line), B2 (red line), and C (green line) municipalities. Panel (b) plots the quarterly number of new listing of newly built homes (normalized to 1 in 2014Q3) in B1 (blue line), B2 (red line), and C (green line) municipalities. Sources: *ECLN* database, an exhaustive record of new home sales transactions in France made available by the French Housing Ministry - see appendix A for a detailed description of the data.



**Figure E.6: The tax responsiveness of residential investment** The figure plots the total quarterly value (in 2019 billions of EUR) of new housing units purchased from developers in France from 1995 to 2019. Dashed vertical green lines indicate the beginning of a more generous investment tax credit or bonus depreciation scheme for new housing; continuous vertical red lines indicate a switch to a less generous tax regime. The successive tax regimes are described in table 4. The total value of units sold is computed from microdata in the *ECLN* database, an exhaustive survey of developer-led projects of five or more units.



**Figure E.7: The tax responsiveness of buy-to-let investment** The figure plots the share of investors (as opposed to owner-occupiers) in the value of new housing units purchased from developers in France from 2010 to 2019. The continuous vertical red line indicates the end of the more generous version of the *Scellier* scheme in 2011, and the switch to a lower level of the tax credit. The dashed vertical green line indicates the beginning of the more generous *Pinel* investment tax credit in October 2014. The successive tax regimes are described in table 4. The share of units sold to investors is computed from microdata in the *ECLN* database, an exhaustive survey of developer-led projects of five or more units.