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Do political parties matter for local land use policies?

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ABSTRACT

Despite interest in the impact of land use regulations on housing construction and housing prices, little is known about the drivers of these policies. Conventional wisdom holds that homeowners have an influence on restrictive local zoning. In this paper, we contend that the party controlling local government might make a major difference. We draw on data from a large sample of Spanish cities for the 2003–2007 political term and employ a regression discontinuity design to document that cities controlled by left-wing parties convert much less land from rural to urban uses than is the case in similar cities controlled by the right. The differences between governments on the two sides of the political spectrum are more pronounced in places with greater population heterogeneity and in those facing higher housing demand. We also present evidence suggesting that these partisan differences might ultimately impact on housing construction and housing price growth.

1. Introduction

Housing construction grew at an extraordinary pace during the last economic boom. In the period 2003–2007 more than 18 million housing units were built in the US, roughly 15% of its historical record (see American Housing Survey, www.census.gov/housing/ahs). In Spain, our case of study, growth was of a similar magnitude, with 4.3 million new housing units being built during the same period, representing 17% of the housing stock. In both cases, such growth markedly increased the area of land under development while reducing overall urban density. For instance, in the US, 80% of the units built were single-family homes. In Spain, the amount of developed land rose by more than 30%, whereas the population grew by just 12% (see www.catastro.es and www.ine.es), gradually changing the landscape to one characterized by low-density sprawl as in many areas of the US.

The acceptance or otherwise of such development varies from one stakeholder to another. Homeowners, it is claimed, dislike development because of its impact on the quality of life in the community and/or on housing values (see, for example, Brueckner and Lai, 1996; Ortalo-Magne and Prat, 2011). Environmentally sensitive citizens worry about the loss of valuable open spaces (European Environmental Agency, 2006) and about the impact of pollution and increased resource consumption (see, for example, Kahn, 2000). Renters and potential new home-buyers welcome the improvement to housing affordability brought about by such developments (Glaeser and Gyourko, 2003). Developers and/or owners of undeveloped land see development as an opportunity to increase their profits (Glaeser et al., 2005a; Hilber and Robert-Nicoud, 2013). The unemployed and employed in the construction and tourism industries see their possibilities of finding or retaining a job enhanced.

Little is known about how governments take into consideration this wide array of interests when determining their land use regulations. Most of the zoning literature holds to the view that it is the homeowners that control the political process (Fischel, 1985, 2001). However, this narrow view is probably a reflection of the almost exclusive focus in the literature on zoning policies in the suburbs of US cities, where the median voter is a homeowner that commutes to work (and who, therefore, sees no job gains from such development), where population is highly homogenous, and where direct democracy regarding such issues is common. Yet, any empirical evidence in favor of this hypothesis is scarce (Dhering et al., 2008), suggesting the need to look elsewhere for a fuller picture. Indeed, various authors have recently provided evidence that interest groups, comprising both developers and environmentalists, might also be fairly influential (Glaeser et al., 2005a; Hilber and Robert-Nicoud, 2013; Solé-Ollé and Viladecans, 2012). The role played by pro-growth coalitions was also highlighted in Molotch's classical study (1976), in which the term 'urban growth machine' was first coined. Fischel (2001, chap. 5) also recognizes the relevance of job creation motives for the zoning policies of rural areas and large cities. In these more heterogeneous communities, the

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role for groups other than homeowners might acquire greater importance, since political parties – known to have preferences regarding land use policies that are more closely in line with those of some of the aforementioned groups – might find it more difficult to commit themselves to the policies desired by the median voter (Ferreira and Gyourko, 2009). Such effects might be further enhanced in places where decisions depend on representative democracy, above all in multiparty systems employing proportional representation.[1] In such situations, party platforms and policies can be more extreme, catering to interests regarding land use regulations that differ from those of the median voter (Schofield, 2007). Thus, eventually, the local land use regulations that are introduced might well depend on the party (or coalition) controlling local government – and, hence, on the social groups that wield most influence over them.

To the best of our knowledge, no previous studies have been undertaken to ascertain the role that political parties play in local land use regulations,[2] albeit that a few do document the relevance of voter ideology for local land use policy (see, for example, Dubin et al., 1992; Gerber and Phillips, 2003). In a recent paper, Khan (2011) reports that the liberal cities of California (i.e., those with a high percentage of voters registered as Democrats, or as supporters of the Green Party or the Peace and Freedom Party) do not grant as many building permits as their non-liberal counterparts. However, it should be borne in mind that these studies do not address exactly the same issue as the one that concerns us here. For instance, the finding that liberal communities impose strict regulations informs us about the preferences of the median voter, but tells us little about the specific influence of a political party. If electoral competition is strong, parties with disparate views in relation to their devising of land use policies might be forced to adapt their platforms to the preferences of the median voter. Were this to be the case, it might be that the party brand does not matter at all in the case of land use policies. Or, should it be found to matter, it could simply be because certain policy drivers (including voter preferences and demand shocks) are correlated with party brand (e.g., left-wing controlled local governments tend to have a higher percentage of left-wing core supporters that have more extreme preferences regarding land use policies).

This makes the identification of the effects of a particular political party on land use policies a far from straightforward task. To tackle the problem we follow a number of recent studies that adopt a regression discontinuity design to identify the effects of political parties on policies (see, for example, Lee et al., 2004; Lee, 2008). Intuitively, the method consists of regressing the outcome variable of interest on a dummy indicating whether a given party won more than 50% of the vote (and therefore holds incumbency) controlling for a flexible function of the vote share. Pettersson-Lidbom (2008), Ferreira and Gyourko (2009) and Gerber and Hopkins (2011) use this methodology to analyze the effect on a broad range of local fiscal policies, although they do not specifically study land use regulations.[3]

Here, we adapt this methodology to the representative democratic system used at the local level in Spain. We have to deal with the fact that many local governments in Spain are coalitions and with the peculiarity of the method used to allocate seats (namely, the 'd'Hondt' rule), which generates many possible thresholds at which one more vote can give a party an additional seat. Specifically, we use the discontinuity at the 50% seat threshold and focus on close elections, defined as those in which the left-wing ideological bloc (i.e., the group of parties classified as left-wing) just won/lost in terms of the number of votes needed to secure a majority of seats in the local council. In justifying this procedure we show that most government coalitions in Spain are formed along ideological lines – i.e., majorities secured by a left-wing ideological bloc tend to generate left-wing controlled governments, defined as those led by a left-wing mayor. This method is then used to estimate the effect of left-wing controlled local governments on the amount of new land assigned for development during a term-of-office, which is the primary land use policy decision that can be taken by Spanish local governments (see also Solé-Ollé and Viladecans, 2012).

Our main results suggest that party brand is relevant. During the 2003–2007 term-of-office, the amount of land assigned for development by a left-wing government was 65% less than that assigned by a similar right-wing government. These partisan differences are most marked in places with greater population heterogeneity and in those facing higher housing demand. We also present evidence suggesting that partisan differences in land use policies ultimately translate into differences in both housing construction and housing price growth. As such, our results are also relevant to the literature studying the effects of land use regulations on various housing outcomes, including housing construction (Cunningham, 2007), housing prices (Glaeser et al., 2005b; Glaeser and Ward, 2006), the size of the housing bubble (Glaeser et al., 2008; Huang and Tang, 2012), urban sprawl (Brueckner and Helsey, 2011) and land use patterns (Konishi, 2013).

The rest of the paper is organized as follows. In the next section we briefly discuss why, and under which particular circumstances, different political parties can be expected to implement different land use policies. In Section 3, we present institutional details about our case study area, Spain: the organization of local government, the system of land use regulation, and the position adopted by Spain's political parties on this issue. The empirical methodology and the data used are outlined in Section 4. Section 5 presents the results and Section 6 concludes.

2. Theoretical discussion: why and when should parties matter?

In this section we discuss the conditions under which we can expect land use policies (in particular, the amount of land converted from rural to urban uses) to depend on party brand. The setting we have in mind is that of a municipality whose citizen-voters have preferences regarding the expansion of the amount of developable land (which we label as Δv) distributed along a line that goes from zero ($\Delta v = 0$) to the maximum value needed to accommodate all future housing projects ($\Delta v = \Delta v^{Max}$). To simplify, we assume that there are two political parties, left-wing vs. right-wing, representing voters that dislike/like growth. The left-wing/right-wing party favors an amount of development that is lower/higher than that preferred by the median voter ($0 < \Delta v^{Left} < \Delta v^{Med} < \Delta v^{Right} < \Delta v^{Max}$). These two parties stand for election on a platform promising to allow a given amount of development during the next term-of-office.

Dating back to Black (1958) and Downs (1957), many formal models of two-party electoral competition have predicted convergence towards the policy desired by the median voter or, more

[1] The influence of specific institutions in determining land use policies has been studied in Lubell et al. (2009) and Gerber and Phillips (2004, 2005).

[2] Many papers have, of course, analyzed the impact of political parties on policies at the federal (e.g., Lee et al., 2004; Lee, 2008) and state levels (e.g., Plotnick and Winters, 1985; Garand, 1988, and Erickson et al., 1989), while just a few recent papers examine their impact on local fiscal policies (see Ferreira and Gyourko, 2009, and Gerber and Hopkins, 2011, for the US, and Pettersson-Lidbom, 2008, and Folke, 2012, for Sweden, and Freier and Odendahl, 2011, for Germany). The US studies conclude that parties do not matter greatly at the local level, while the European studies report a more relevant role at this level for political parties.

[3] Regression discontinuity design has been used to evaluate a wide range of policies (see Imbens and Lemieux, 2008, and Lee and Lemieux, 2010, for recent surveys). Recently, it has also been applied to evaluate the effects of land use policies (see Turner et al., 2011; Cyrus et al., 2011).

generally, towards the center of the political spectrum. In our context, these models would suggest that the amount of land made available for development by the left- and right-wing parties (which we label as Δu) would be the same and equal to the amount preferred by the median voter ($\Delta u^{Left} = \Delta \upsilon^{Med} = \Delta u^{Right}$). As this prediction has been contested by many empirical studies (for policies other than land-use regulations), recent theoretical work has tried to reconcile these findings. First, Alesina (1988) and Besley and Coate (1997) suggest that the lack of credibility of campaign promises accounts for the discrepancies between a party's platform and the policies it subsequently implements. Second, strategic extremism might also generate divergent policies (e.g., Glaeser et al., 2005c) with a party proposing more extreme platforms in order to obtain more voters among its core supporters, either through an increase in turnout or through resource mobilization. These models would therefore predict that the amount of land made available for development would lie somewhere between the amount preferred by the party and that preferred by the median voter ($\Delta \upsilon^{Left} < \Delta u^{Left} < \Delta \upsilon^{Med}$ and $\Delta \upsilon^{Med} < \Delta u^{Right} < \Delta \upsilon^{Right}$). Similar or perhaps even more markedly divergent policy platforms could result in the context of multi-party elections with voters caring about the quality of candidates (e.g., Schofield, 2007).[4] This is the institutional setting that most resembles the Spanish case.

Some authors are skeptical about the relevance of the policy-divergence prediction at the local level. First, in line with Tiebout (1956), individuals could be assumed to choose their municipality of residence according to their preferences for local public goods, at least within a local labor market. The outcome of this process would be a sorting of individuals into more homogeneous communities. Then, with lower intra-municipal demand heterogeneity, political discrepancies should be much smaller. Similarly, with less heterogeneity, the promises of politicians should be more credible and the ability to target core supporters with extreme preferences should be lower (see, for example, Ferreira and Gyourko, 2009). The relevance of this line of reasoning might be limited by the (relatively) low degree of residential mobility in Spain, by the fact that the majority of people in certain areas lives and works in the same place, and by the substantial intra-city heterogeneity in our sample.

Second, it could also be argued that decisions related to the provision of local public services are of a largely technical nature and do not involve policy preferences (i.e., 'there is no right- or left-wing way of picking up garbage') and that policy differences are more likely to be found in areas related to redistribution or to moral issues, which are the responsibility of higher tiers of government (see Gerber and Hopkins, 2011). Moreover, the claim might be made that local land use policies are not a partisan issue, since with proper side payments the policy would provide benefits for all social groups (see Fischel, 1985). However, once again, heterogeneity may hinder the achievement of such deals. Informal evidence suggests that in Spain there is a great deal of ideological controversy over these policies (see next section). Thus, in these two cases, the party brand would not matter simply because there are no discrepancies in the desired amount of development ($\Delta \upsilon^{Left} = \Delta \upsilon^{Med} = \Delta \upsilon^{Right}$).

Third, any partisan discrepancy between land use policies in a specific municipality might depend on whether the issue acquires relevance during the electoral campaign. Unlike fiscal matters, which are always important, the salience of local land use policies and, especially, the amendment of land use plans to allow for more development depends on the situation of the housing market. Consider, for example, the case of a municipality undergoing a substantial demand shock, with the possibility that the amount of land made available for development will not be enough to accommodate the portfolio of existing housing projects. In this case, the platforms of the different parties *could* be very different, the left-wing party opposing and the right favoring the amendment of the plan in order to convert more land for development. However, in a municipality with low housing demand (i.e., when $\Delta \upsilon^{Max} < \Delta \upsilon^{Right}$), the right-wing party will have to adhere to a policy platform that proposes much less development than it would have preferred in an unconstrained situation. This would move the platform of the right-wing party to the center, thereby attracting many votes, while forcing the left-wing party to converge to the center too. High housing demand can thus be expected to lead to a higher degree of policy divergence.

This revision of the aforementioned theories generates expectations as to the possible partisan differences that might appear in relation to Spain's local land use policies. First, there are expectations that partisan differences might be stronger in Spain than in the US, given the country's electoral institutions and the lower degree of residential mobility. Second, in municipalities located in fragmented local labor markets and/or with lower population heterogeneity, partisan differences should be smaller. Finally, differences should be greater in periods and/or areas undergoing strong housing demand shocks.

3. Institutional setting

3.1. Spain's local government

Municipalities are the main tier of local government in Spain, there being more than 8000 local government authorities. Since 1979, the members of these municipal councils have been elected. Elections are now held every four years simultaneously in all the municipalities. Voters choose between various party lists, which being closed means that no preferences regarding the ranking of the names on these lists can be expressed. The electoral system is proportional and seats are allocated according to the 'd'Hondt' rule (more details in Section 4.1). In most municipalities, several right- and left-wing parties run separately, with pre-election coalitions being very rare. Some of these parties adopt more central platforms while others are more extreme, particularly in the case of land use regulations (see Section 3.3 for details). Most of the candidates run under regional or national party brands. There are also many local parties, but they win the mayoralty in just a handful of cases.

The mayor is then elected by a majority of the council (see Colomer, 1995). A substantial proportion of governments are coalitions (around 30% during the term we analyze here), formed most of the times along ideological lines. This rule is not always respected, especially in small towns, where other considerations might matter more than ideological proximity. The council operates as a small representative democracy, and has to reach a majority vote to pass the initiatives and regulations proposed by the mayor, who acts as the agenda-setter. The discipline enforced by Spain's party system means that the chances of amending the mayor's proposals are quite low for mayors controlling a majority of the seats.

3.2. Local land use policies

Land use regulations in Spain are controlled by a very detailed, rigid system (Riera et al., 1991), although they do not differ greatly

[4] In these models, platform divergence can occur whenever there are important centrifugal forces. These forces strengthen as voters' preferred policies become more heterogeneous and the differences in candidate quality become more marked. These models predict that party positions during the campaign can disseminate along a principal policy axis. Post-election coalition bargaining between the parties determines the final policy implemented, lying at some point between the positions of the parties forming the coalition. Several empirical analyses corroborate this theory, especially for proportional electoral laws (see, e.g., Schofield and Sened, 2006).

from the zoning regulations operating in various parts of the US. Land use planning in Spain is essentially a municipal responsibility. There are more than 8000 municipalities, so the system is highly fragmented. Municipalities draw up a 'General Plan', providing a three-way land classification: *built-up land*, *developable land* (the areas of the community where future development is allowed), and *non-developable land* (the rest of the territory – agrarian and other uses, where the development process is strictly prohibited, at least until a new plan is approved). In theory, the 'General Plan' has to be updated every eight years, but the land classification can be quite easily modified before that date. The amendment plan, known as a 'Partial Plan', is also a legally binding document. The amount of *developable land* can be considered the main land use policy instrument available to the local government, and this is the variable we analyze in this paper. Of course, the plan includes very detailed regulations regarding many other aspects: land zoning (residential, commercial, industrial), floor-to-area ratios, setting aside of land for streets, green spaces and public facilities, etc. While it would be of great interest to analyze these other regulatory dimensions, no data are available to measure them.

3.3. Political parties and land use

Most members of Spain's local councils are elected from national or regional party lists. After the 2003 elections, the two main national parties (i.e., the left-wing 'Partido Socialista Obrero Español', PSOE, and the right-wing 'Partido Popular', PP) held 71% of the mayoral offices (i.e., 36% of the offices were occupied by the PSOE and 35% by the PP). The remaining mayoralties were distributed as follows: 10% were held by other left-wing parties,[5] 13% by various regionally-based right wing parties, and 6% by local parties.

The parties on the left and right of the political spectrum hold very different views as to how land use policies should be designed. These differences can be documented by looking at the national or regional manifestos these parties launched before the local elections. In recent elections, much attention has been devoted to environmental and other problems caused by excessive growth and urban sprawl. The proposals included in the manifestos of the main national left-wing parties (PSOE and IU) are illustrative of the emphasis placed on restricting urban growth. For example, the PSOE program proposes[6]:

> "To establish limits to urban growth (...) based on the real and potential economic and demographic demand in the city, on the capacity to absorb growth, on the stock and capacity of existing infrastructure, and on the natural environment."

The program of the former communist party, IU, likewise includes a number of proposals related to land use policies, concerned primarily with the protection of the environment[7]:

> "To promote a compact city as opposed to a diffuse city. To preserve non-developable land in order to protect the environment. To create green belts surrounding the city, combining parks and agricultural lands."

The programs of these two parties also include several proposals for dealing with the housing affordability crisis; yet, neither proposes making housing more affordable by allowing more land to be developed. The PSOE proposed reserving 25% of all developed land for social housing while IU proposed the public provision of rental social housing. The programs of the other left-wing parties include similar proposals, lying somewhere between those of the PSOE and IU. In general, it could be said that that the discourse of Spain's left-wing parties tend to vilify urban growth while ruling out any relationship between an increased provision of urban land (or the easing of regulatory constraints, in general) and housing affordability. Thus, it seems quite reasonable to expect a tendency among these parties to restrict growth in the amount of *developable land*, and for parties located to the far left (i.e., IU) to enact even more restrictive land use policies.

This is in marked contrast with the position taken by right-wing parties and, especially, with that adopted by the PP. The local manifesto of the country's leading right-wing party does not include any specific proposals related to the containment of urban growth. It does, however, include some proposals related to easing regulatory constraints, namely "to improve and simplify the process of urban development; and to speed up the completion of urban development projects".[8] The discourse of the PP is the one that places greatest emphasis on the virtues of deregulation of the land market as a means of improving housing affordability. Some of the other right-wing parties also adhere to this view, while others are more moderate, but they are generally in favor of urban planning in order to minimize the adverse impacts of growth.[9] Politically, therefore, this group can be considered as lying somewhere between the PP and the PSOE.

4. Empirical analysis

4.1. Empirical design

OLS *with controls*. As a first approach, we estimate the effect of left-wing governments on land policies using OLS, controlling for a set of covariates and including area fixed effects:

$$\Delta u_{ij} = \alpha\ dLeft_i + X_i'\beta + f_j + \varepsilon_{ij} \tag{1}$$

where Δu is the increase in the amount of land placed under development during the term-of-office (i.e. the amount of land converted from rural to urban uses) in municipality i located in area j. The dummy *dLeft* is equal to one in the case of a left-wing government and zero in the case of a right-wing government. The vector X includes control variables measuring influences on local land use decisions, related either to the intensity of the housing demand shock experienced by each municipality during the period or to the preferences of the resident population for (or against) growth. We describe these variables in detail in the next section. f_j are local area fixed effects, one for each of the urban areas identified and also one for each of the rural sections in each Spanish province. These fixed effects control for any omitted influences on land policies (e.g., economic cycle, area-wide amenities) that are common to the municipalities located in the same local area.

One advantage of this approach over previous methods reported in the literature (see, for example, Bates and Santerre, 1994, 2001; Evenson and Wheaton, 2003) is that the dependent variable can be precisely matched to the particular government responsible for the policy at that time. Its drawback, however, is the possibility that certain influences on land policy that are correlated with the partisan identity of the local government remain omitted. For instance, it might well be the case that pro-growth residents are concentrated in certain municipalities of the urban area and so tend to vote for right-leaning parties. It might also

[5] On the left, the other main party is 'Izquierda Unida', IU, but there are also some minor left-wing regional parties. On the right, the most important party is 'Convergencia i Unió', CiU, in Catalonia.

[6] PSOE (2006): "Para una nueva política urbanística y del territorio", Programa Marco Elecciones Municipales 2007. http://www.psoe.es/organizacion/docs/454856/page/programa-marco-elecciones-2007.html.

[7] http://izquierdaa-unida.es/sites/default/files/doc/Programa_Marco_Municipal.pdf.

[8] http://www.elpais.com/elpaismedia/ultimahora/media/201105/05/espana/elpepunac_7_Pes_PDF.pdf.

[9] See, e.g. the manifesto of the main right-wing party in Catalonia, CiU, http://ciu.cat/media/55510.pdf.

be the case that places affected by municipality-specific demand shocks during the period analyzed turn to the right in order to facilitate the development projects being implemented. In both instances, failure to account appropriately for the residents' ideology (or for the intensity of the housing demand shock) would bias the *dLeft* coefficient.

Regression discontinuity. To deal with the omitted variables problem a number of papers have recently adopted the close-race regression discontinuity (RD) design framework (e.g. Lee et al., 2004; Lee, 2008; Pettersson-Lidbom, 2008; Ferreira and Gyourko, 2009; Albouy, 2013; Folke, 2012; Gerber and Hopkins, 2011). The idea underpinning this methodology is that elections won by a given party by a narrow margin are very similar to the elections lost by that party by a narrow margin. Thus, by focusing on close-races, the RD design generates quasi-experimental estimates of the effects of interest (see Hahn et al., 2001). In a recent survey, Green et al. (2009) show that RD designs are comparable in their accuracy to experimental studies.

However, the application of this methodology is not straightforward in our case. In Spain, the proportional representation system used at local elections means that it is less evident that the partisan control of the government changes at a given vote threshold. Firstly, the rule used to allocate seats generates many possible thresholds at which an additional vote can bring a party one more seat. Briefly, for each party obtaining more than 5% of the vote, the d'Hondt rule computes a series of 'comparison numbers' by successively dividing its votes by 1, 2, 3, 4, etc. The 'comparison numbers' of all parties are then ranked and a given number of seats are allocated to the parties on the basis of this ranking (see the Online Appendix for an example illustrating the application of the d'Hondt rule). For each party's marginal seat, there are an additional number of votes that need to be won in order to gain an extra seat (or which must not be lost in order to hold onto this seat). As such, each party and each seat has a specific vote threshold. Secondly, in a non-trivial proportion of municipalities no party has more than 50% of the seats, the mayor being elected by a coalition of parties. There is thus no straight relationship between the number of seats held by a party or group of parties and their control of local government.

To deal with these difficulties we proceed in two steps. Firstly, we build on the fact that, in Spain, a high proportion of local government coalitions are formed along ideological lines: majorities of seats held by left-wing parties tend to generate left-wing controlled governments. This allows us to use the discontinuity at the 50% seat threshold, and so to consider as close elections those in which the left-wing ideological bloc has won/lost by just one seat. By so doing, we are comparing two potential ideologically connected coalitions (i.e., left-wing vs. right-wing) with a seat difference of just one seat. For this procedure to be appropriate, ideology must be a powerful driver of coalition formation at the local level. Admittedly, however, other idiosyncratic factors at the local level (e.g., personal relationships, historical disagreements, the need to replace a bad incumbent, etc.) might carry sufficient weight to impede the formation of an ideologically connected coalition. Yet, anecdotal evidence seems to support the claim that in Spain ideology is a very powerful determinant of coalition formation at the local level. As explained, most of the candidates in Spain run under regional or national party brands, so there is a tendency for ideological coalitions formed at the regional level to be reproduced at the lower level. Moreover, this also seems to be backed up empirically, since the fact of one ideological bloc of parties holding a majority of seats is a very strong predictor of the ideological stance of the mayor (see Section 5.1).

This means that any departures from the ideological motive for coalition formation can be handled by our empirical methodology. We use a 'fuzzy' RD design, which allows the jump in the probability of having a left-wing government at the 50% seat threshold to be lower than one (see Van der Klaauw, 2002; Lee and Lemieux, 2010). Since the probability of treatment jumps by less than one at the threshold, the jump in the outcome variable (e.g., Δu) at this point can no longer be interpreted as an average treatment effect. However, the treatment effect can be recovered either by dividing the jump in the outcome variable by the jump in the probability of treatment or by estimating the effect of party control by 2SLS, using the threshold dummy as an instrument for party control. As in any 2SLS, it is crucial to have a powerful first-stage – corresponding here to the fact that having a majority of left-wing seats is a good predictor of there being a left-wing mayor. As discussed above, we will show that this is the case.

Secondly, note that elections which are close in terms of seats are not necessarily that close in terms of number of votes. The ideological bloc holding the majority of seats might have won this last seat by just a few votes or by many and have just missed out on winning an extra seat. This means that we need to take into account how many votes the party holding this marginal seat would have to lose if it were to lose the seat. In our context, close-elections are precisely those in which just a few votes are needed to move this marginal seat from one bloc to the other. Whether the marginal party won one vote more or less can be considered a random event, and this is why municipalities located close to each side of this threshold can be treated as being similar. The difference with a standard close-election RDD is that the calculation of this vote distance is not straightforward. In this paper, we develop a method for computing this vote distance which takes into account the specificities of the 'd'Hondt' rule. The details of the method and the assumptions underlying its calculation are provided in Section 4.3 and in the Online Appendix. This vote distance variable is then used as the forcing variable in our RDD analysis. Thus, instead of controlling for the non-linear distance in seats to the seat majority threshold, we are able to control for the distance in votes to the seat majority.

Once this distance has been computed, the reduced-form equation used to estimate the effect of party identity on local land supply can be expressed as:

$$\Delta u_i = \lambda d(\textit{Left seats} > \textit{Right seats})_i + f(\%\ \textit{Votes to left-wing majority})_i + \upsilon_i \tag{2}$$

where d(*Left seats* > *Right seats*) is a dummy equal to one if the left-wing bloc has more seats than its right-wing counterpart and, thus, defines the threshold, and f(% *Votes to left-wing majority*) is a non-linear function (e.g., a polynomial or a locally weighted regression) of the distance in votes to the change to a left-wing bloc seat majority, fitted separately to both sides of the threshold. Alternatively, the following equation could be estimated by 2SLS:

$$\Delta u_i = \delta\ dLeft_i + g(\%\ \textit{Votes to left-wing majority})_i + \xi_i \tag{3}$$

using d(*Left seats* > *Right seats*) as the instrument for *dLeft*. The δ coefficient is a 'Local Average Treatment Effect' (LATE). The first-stage equation is as follows:

$$dLeft_i = \gamma d(\textit{Left seats} > \textit{Right seats})_i + h(\%\ \textit{Votes to left-wing majority})_i + \omega_i \tag{4}$$

where $g(\bullet)$ and $h(\bullet)$ are also non-linear functions of the distance in votes to seat majority. If the order of the polynomials used is the same, then the LATE can also be obtained as the ratio between the reduced form coefficient and the first-stage discontinuity (i.e., $\delta = \lambda/\gamma$).

4.2. Econometrics

The estimation of the OLS equation with controls is straightforward. The estimation of the RD equation with close elections requires taking various methodological decisions into account. First, our main estimates use the whole sample and controls for a flexible polynomial. We explicitly test for the optimal order of the polynomial using the Akaike information criteria. This method has the advantage of using all the observations and, thus, of improving the efficiency of the estimator. However, by not restricting the bandwidth to a vicinity of the threshold we run the risk that some extreme observations may have an influence on the estimated effect. In our case, moreover, there is an additional problem. As we show in the next section, besides the vote discontinuity that determines that gaining the last seat gives a majority, there are also the discontinuities that determine the allocation of the infra-marginal seats. By using the whole sample, the estimated polynomial relies on information that overlaps with the areas surrounding these other discontinuities. We consider this not to be an excessively grave problem since, as we show below, the increase in the number of seats below the one which finally gives the majority of seats has a very small impact on the probability of controlling government. Despite this, we also present results for a restricted bandwidth. The optimal bandwidth – computed following the procedure proposed in Imbens and Kalyanaraman (2009) – was found to be around 25%. So, following the recommendation made by Lee and Lemieux (2010), we also present the results for the optimal and half optimal bandwidth, using in this case a locally weighted regression. The half optimal bandwidth is somewhat smaller than the maximum vote distance for the sample of close elections (i.e., where the distance to seat majority in terms of seats is either -1 or $+1$). This constitutes, therefore, a way of checking that our results are not influenced by the use of a bandwidth that overlaps with other (minor) discontinuities.

Second, in order to show that there is a valid case for the RD design proposed, we verify the discontinuity in the probability of treatment. We examine the discontinuity graphically and we estimate the jump in the probability of treatment using the whole sample and a flexible polynomial and the reduced bandwidths with a locally weighted regression. Third, we also check the continuity of the forcing variable around the threshold by looking at the histogram, as well as by using a more formal test (see McCrary, 2008). The continuity test is a means of discarding the manipulation of the forcing variable, a problem that some authors suggest can occur in close-election RD designs (Caughey and Sekhon, 2011). With the same purpose in mind, we also test for the continuity of the pre-determined covariates. Finally, we present the results both without controls and controlling for the same covariates as those used in the OLS analysis.

4.3. Data

Sample. We carried out our main analysis using data from a sample of 2112 Spanish municipalities for the 2003–07 term-of-office. These years coincided with the peak in the last housing boom, a period in which the conflict between pro- and anti-growth groups was particularly intense and, hence, the perfect setting for the testing of our hypothesis. Although our land use data are available on a yearly basis, we decided to use a long time difference. The dependent variable is, therefore, the increase in developable land between 2003 and 2007, and the control variables refer to the beginning of the period. There are several reasons for this choice. First, political variables (e.g., *dLeft*) can only be measured once, which is when an election takes place. This means that there is no real statistical gain to be made in using yearly data. Second, the dependent variable does not change every year; developable land only changes when a new urban plan is passed, and this is a fairly rare occurrence, happening more frequently when the real estate sector is booming. Thus, by aggregating the data over the term we considerably reduce the number of censored observations in our sample. This helps to reduce data volatility, which is crucial for improving the efficiency of the estimates.

The eventual sample of 2112 municipalities reflects the availability of our data. Spain has about 8000 municipalities, but most of them are small (i.e., 90% have fewer than 1000 residents). The database providing information on land use categories covers the whole of Spain, but some of the other databases used are restricted to municipalities with over 1000 inhabitants, which means that the smallest municipalities have been eliminated from our sample. We have also eliminated from our sample those municipalities for which we either lacked political data or for which the data were not reliable. We believe the final sample to be representative of the whole population.[10] Subsequently, because of a lack of data, we also use the subsample of 252 municipalities with more than 25,000 inhabitants. This subsample obviously differs in many dimensions from the whole population and, as such, we are unable to claim that the results can be generalized.

Land policies. The data used to measure the amount of developable land are taken from the Spanish property assessment agency (*Dirección General del Catastro*) and are derived as a by-product of the assessment process that this agency undertakes on all properties in the country. Although the values of properties are only reassessed from time to time, the up-date in the traits of each property (and so its classification as developed, developable but vacant, or non-developable) is conducted yearly. This is the only statistical source of data covering the whole of Spain that can be used to measure the land use category of undeveloped land plots. Note that GIS data (e.g., coming from the *Corine Land Cover* project, Ministerio de Fomento, 2006) do not help much in this respect, because they only measure what can be seen (already developed land) not what has been approved by the local government but does not yet physically exist (land which may be developed).

Likewise we present some results using other dependent variables, namely, the growth in developed land and the change in housing prices over the same period. Data on developed land are also drawn from the Spanish property assessment agency, and are available for the same sample of municipalities. Data on housing prices are provided by the Spanish government (Ministerio de Fomento) and come from private assessment firms. The main drawback here is that the government only discloses information for municipalities with more than 25,000 residents. A further problem is that the information provided is simply the average price per m^2 for all transactions, which makes it impossible to consider any degree of heterogeneity. As such, the results for the housing prices need to be treated with considerable caution.

Party classification. We have information on the number of votes and seats obtained by each party at the 2003 local elections. We also know the party identity of the mayor during the 2003–2007 political term. We classified the parties in five groups: Left–Left, Center–Left, Center–Right, Right–Right and Local parties. Based on informal evidence regarding the position adopted by each party on matters relating to land use regulations (see Section 3.3), we

[10] In the Online Appendix we provide descriptive statistics for the whole population of Spanish municipalities and for different subsamples (municipalities larger than 1000 residents, larger than 1000 and with all information available). The restricted sample with more than 1000 inhabitants is eventually very similar to the unrestricted sample.

classified the main left-wing party (PSOE) as Center–Left and the main right-wing one (PP) as Right–Right. The former communist party (IU) was classified as Left–Left; also in this group we included many small or even local extreme left-wing and green parties and some of the left-wing regional parties. The Center–Right group includes the right-wing regional parties. Local independent parties were either included in the Center–Right group or excluded from the analysis on the grounds that, while local left-wing parties tend to identify themselves as such (by choosing such labels as 'green' or 'progressive'), the other local independents are probably centrist or right-wing. In any case, note that we have just 78 observations (from a total of 2112) of mayors representing Local parties and the results are unaffected by their exclusion.

Overall, the proportions of municipalities allocated to the four groups are 6.7%, 44.3%, 14.9% and 33.9%, for Left–Left, Center–Left, Center–Right and Right–Right, respectively. If we consider just the close-election sample (one seat from a majority) the proportions are more or less the same: 3.6%, 42.5%, 16.2% and 37.5%, respectively. The *dLeft* dummy is equal to one for mayors from the parties in the Left–Left and Centre–Left groups. The *d*(*Left seats* > *Right seats*) is equal to one for municipalities where the seats from parties in the first two groups are higher than those from the last two groups. We also use this information to obtain the results when restricting the sample to pairs of ideological groups: Center–Left vs. Center–Right, Center–Left vs. Right–Right, Left–Left vs. Center–Right and Left–Left vs. Right–Right. Although the classification of parties into these groups might seem somewhat ad hoc, it should be stressed that the results are robust to the displacement of the minor parties to adjacent groups (results available upon request). The reason for this is that each of the groups is dominated by one of the main parties. Moreover, below we provide results comparing the municipalities controlled by one or other of the two leading national parties (PP vs. PSOE). These results are very similar to those that use the broad left vs. right categories.

Vote distance measures. To compute the distance in votes to a change in a majority of the seats (% *Votes to left-wing majority*), we develop an algebraic formulation for the 'd'Hondt' rule, the system used in Spain's local elections to translate votes into seats. The easiest way to perform this calculation is simply to take the cases in which the incumbent's ideological bloc won by one seat, identify the party holding the seat and then count how many votes would have to be taken away from this party for it to lose this seat. Under the 'd'Hondt' rule, the formula for this calculation is relatively straightforward. To simplify the problem, let's assume that the party assigned such a seat forms part of the same ideological bloc as that of the incumbent and that the party competing for this seat (i.e., the party that would have won this seat had the other party lost a certain number of votes) belongs to the ideological bloc of the opposition. In this case, the number of votes needed to lose the marginal seat is simply the difference between the 'comparison number' for the last seat won by the party in the incumbent's bloc and the 'comparison number' for the next seat to be won by the party in the opposition's bloc. This method for computing the vote distance is based on a number of implicit assumptions. Note, for instance, this procedure assumes that the votes taken away from the party holding the marginal seat are transferred only to the abstention vote. It also assumes that shocks affecting one party are independent of shocks affecting other parties in the same bloc. Clearly, there is no precise procedure for dealing with these questions, as it requires information about the migration of votes to abstentions and/or to parties in the other bloc, and about the co-movement between shocks affecting parties in the same bloc. This situation is further complicated in the case of local elections, given the diversity of situations. Here, we opt for a more feasible approach and compute this vote distance under different assumptions. We then seek to verify whether the results are robust to the vote distance measure used.[11]

In our preferred vote distance measure – for which we present our main results – we assume that the votes taken away from the party holding the marginal seat are transferred only to abstentions and not to the parties in the other bloc.[12] We also assume that negative vote shocks simultaneously affect all the parties within the incumbent's ideological bloc,[13,14] so we subtract votes not just from the party holding the marginal seat but from all parties in the bloc in proportion to the initial votes received by each party. Intuitively, our method works as if we were subtracting small numbers of votes from one of the blocs, distributing these votes between the parties of that bloc according to their vote share, while keeping the number of votes for the parties of the other bloc constant. As we subtract more votes, seats start shifting from one bloc to the other. We stop transferring votes when we observe a shift in seat majority from one bloc to the other (i.e., when the last seat giving the majority to one bloc moves to the other bloc). The number of votes needed to reach this stage, divided by the total number of votes, is our measure of vote distance.[15]

As a robustness check, we use other vote measures that make alternative assumptions regarding vote migration: all votes lost are transferred to the other bloc (and these votes are distributed to the parties in the other bloc in proportion to their initial vote share within the bloc), and the votes lost are transferred to abstentions and to the other bloc.[16] The results are not affected at all by this decision, suggesting that the specific vote distance measure used does not matter.

Control variables. We use the following control variables (data sources provided in Table 1). Firstly, the amount of land assigned for development that remains vacant at the beginning of the period as a proportion of the previous built-up land (% *Vacant Land*). The argument here is that if a lot of land assigned for development remains undeveloped, there will be no immediate need to alter regulations assigning more land for development. Similarly, if there is no vacant land at all, there will be considerable pressure to release more land for development in order to accommodate possible future demands. Secondly, the amount of open land at the beginning of the period as a proportion of previous built-up land (*Open Land*), i.e. the land under the jurisdiction of the municipality which was neither build on nor assigned for development but vacant. If there is a shortage of open land –either because the town grew a lot in the past, or it has a small jurisdiction – the government might opt to preserve scarce open space or postpone development decisions until a later date.

[11] A number of papers employ a regression discontinuity design with proportional election systems (see, for instance, Folke (2012) for Sweden and Freier and Odendahl (2011) for Germany). These papers face the same kinds of difficulties as we face here and their computations of the forcing variable also involve making certain specific assumptions.

[12] We believe this assumption to be plausible in Spain given the importance of vote transfers from/to abstentions. This can be documented by examining the correlation between turnout and the left-wing share of the vote. Using district-level national elections data, Lago (2010) reports a 0.5 correlation between the increase in turnout between two consecutive elections and the increase in the socialist (i.e. PSOE) vote share. Using our municipal-level data we find roughly the same correlation.

[13] The vote outcomes of the two main left-wing groups of parties are highly correlated. Using our municipal-level data we find a statistically significant correlation of 0.37 between the increase in the socialist vote share (PSOE) and the increase in the vote share of more extreme left-wing parties.

[14] This assumption is irrelevant in the right-wing bloc, since there is usually only one dominant party.

[15] In the Online Appendix we provide a numerical example to explain how this procedure works. The Online Appendix also shows the development of the algebraic formulation used in the calculations.

[16] The Online Appendix provides some robustness checks using these other vote distance measures.

Table 1
Definitions and sources of the variables.

	Definition	Sources
%Growth in developable land	[(Built-up land + Vacant land, end of term) – (Built-up + Vacant land, beginning of term)/Built-up land, beginning of term]	DCG, Dirección General del Catastro (2007): "Estadísticas sobre ordenanzas fiscales del Impuesto sobre Bienes Inmuebles", http://www.catastro.meh.es/esp/estadisticas1.asp#menu1. (Built-up land = 'superficie edificada', Vacant land = 'superficie de solares')
%Growth in developed land	[(Built-up land – Built-up beginning of term)/Built-up land, beginning of term]	
%Vacant land	[Vacant land, beginning of term)/Built-up land, beginning of term]	
%Open Land	[Total land area of the municipality – Built-up land beginning of term/Built-up land, beg. term]	INE (www.ine.es) & DCG, Dirección General del Catastro (2007)
%Growth in housing prices	Average annual growth rate in price per m^2 of old and new housing units during term	"Precios de vivienda libre en los municipios de más de 25,000 habitantes", Ministerio de Fomento, http://www.fomento.gob.es/
dLeft	Dummy = 1 if the mayor belongs to a party classified as left-wing	Ministerio del Interior, Base Histórica de Resultados Electorales, http://www.elecciones.mir.es/MIR/jsp/resultadosindex.htm. & El País (1999, 2003): 'Anuario Estadístico'
d(Left > Right)	Dummy = 1 if the parties classified as left-wing have more seats in the local council than those classified as right-wing	
%Votes to left-wing majority	% of Votes needed by the left-wing bloc to either lose the last seat they hold or to win an additional seat (see Online Appendix for a details of the method used in the computation)	
dUrban	Dummy = 1 if municipality belongs to an urban area	AUDES project: 109 urban areas defined using aerial photographs on the basis of geographical continuity (see www.audes.es),
dSuburb	Dummy = 1 if municipality belongs to an urban area but it is not the central city	
Amenity index	[Houses with problems related to: noise, dirt, crime, pollution, or lack of green space, as of 2001/Houses in 2001]	INE (www.ine.es), 2001 Census of Buildings
Road accessibility	[Houses with poor accessibility to roads, as of 2001/Houses in 2001]	
%Aged 25–40	[Residents aged 25 to 40 beginning of term/ Resident population beginning of term]	INE (www.ine.es), 2001 Census of Population & 'Estadística de Variaciones Residenciales' (several years)
%Immigrants	[Immigrants arrived during the term/Resident population beginning of term]	La Caixa (2001): 'Anuario Económico de España'
%Out-commuters	[Commuters in 2001/Resident population in 2001]	
%Homeowners	[Houses occupied by owner in 2001/Houses in 2001]	
%Graduate	[Residents with a higher education degree in 2001/Resident population in 2001]	
%Unemployed	[Residents which were unemployed, beginning of term/Resident population, beginning of term]	
Population size	Resident population, beginning of term	
Income per capita	Personal income, beginning of term/Resident population, beginning of term.	

Thirdly, a basic set of control variables *Z*, measuring the main traits that account for recent urban growth in Spain, and which includes the *Urban*, *Suburb* and *Beach* dummies. The European Environmental Agency, 2006 notes that most of the recent housing growth in Spain has been concentrated in these places, so we expect them to capture a large share of the spatial variation in the demand for land. Fourthly, a full set of local area dummies f_j. These effects are included because the size of the increase in demand depends to a great extent on certain geographical traits (e.g., weather, proximity to the coast, regional regulatory framework, and major infrastructure such as ports or airports) that are common to municipalities located near one another. We use 109 urban area and 50 provincial dummies.[17,18]

Finally, we also use a set of additional control variables, *W*, measuring either the size of the demand increase or the pro- or anti-growth preferences of the residents. This set includes: (a) Exogenous measures of local demographic shocks: % *Aged* 25–40, which measures the number of potential new families at the beginning of the period, % *Immigrants* (i.e. those that arrived during the period, expressed as % of residents at the beginning of the period); (b) variables that account for the amenity and productivity factors deemed important for location decisions (i.e., an *Amenity index* and a measure of *Road accessibility*); (c) variables more closely related to a resident's preferences for development, but also arguably correlated to 'demand pressures' (i.e. % *Out-commuters*, % *Homeowners*, % *Unemployed*, % *Graduate*, *Population size*, *Density* and *Income per capita*).

5. Results

5.1. OLS with controls

Table 2 presents the results of the estimation of equation 1 by OLS. Column (i) presents the results without controls. Column (ii)

[17] Since both sets of dummies are introduced simultaneously, the provincial dummies account for the effects common to all municipalities in the non-urban portion of a province.

[18] The urban areas are those identified by the AUDES project using geographic contiguity criteria (see www.audes.es). Alternatively, we could have used local labor markets (LLM) as defined using commuting patterns (see Boix and Galletto, 2006). The drawback of using this definition of local area is that outside urban areas the number of municipalities per labor market is very low, meaning that in our restricted sample we will have many areas with just one observation.

Table 2
OLS results.

	(i)	(ii)	(iii)	(iv)
dLeft	−0.121 (0.044)***	−0.146 (0.045)***	−0.175 (0.067)***	−0.171 (0.085)**
%Vacant land	–	−0.632 (0.133)***	−0.655 (0.137)***	−0.674 (0.159)***
%Open land	–	0.075 (0.014)***	0.076 (0.013)***	0.079 (0.011)***
dUrban	–	0.081 (0.039)**	–	–
dSuburb	–	0.092 (0.041)***	0.163 (0.070)***	0.091 (0.035)***
dBeach	–	0.134 (0.050)***	0.126 (0.047)***	0.113 (0.046)**
Adj-R^2	0.043	0.092	0.148	0.142
F-est. (*all var.*)	7.33	19.57	4.96	3.93
[*p-value*]	[0.001]	[0.000]	[0.000]	[0.000]
F-est. (*main controls*)	–	23.09	21.32	16.44
[*p-value*]		[0.000]	[0.000]	[0.000]
F-est. (*area effects*)	–	–	4.44	3.09
[*p-value*]			[0.000]	[0.000]
F-est. (*additional controls*)	–	–	–	0.30
[*p-value*]				[0.112]
Main controls	NO	YES	YES	YES
Area effects	NO	NO	YES	YES
Additional controls	NO	NO	NO	YES
Num. Obs.	2112	2112	2112	2112

Notes: (1) Dependent variable: Δu, % increase in developable land over the term. (2) Robust standard errors in parenthesis, *p*-values in brackets. (3) Additional controls: *% Aged* 25–40, *% Immigrants*, *Amenity index*, *Road accessibility*, *% Out-commuters*, *% Homeowners*, *% Unemployed*, *% Graduate*, *Population size*, *Density* and *Income per capita.* (4) Area effects: dummies for each of the 109 AUDES urban areas and for each of Spain's 50 provinces.
* Statistically significant at the 90% level.
** Statistically significant at the 95% level.
*** Statistically significant at the 99% level.

introduces the main set of controls (i.e., the amount of vacant and open land, and the dummies identifying whether the municipality is located in an urban area, whether it constitutes a suburb or it is on the coast). Column (iii) introduces the full set of local area dummies, and Column (iv) controls for a large set of additional covariates. The results indicate that left-wing governments convert less land from urban to rural uses than is the case with right-wing governments. The effect increases as the different sets of controls are added, but it is qualitatively the same in each case. The results of Column (iii), our preferred specification, indicate that the new land that was allowed to be developed during the term (as a proportion of the built-up area at the beginning of the term) is 0.175 less under a left-wing government. That the average value of this variable for the municipalities controlled by the right is approximately 0.55 means that, on average, left-wing governments develop 32% less land than that developed by right-wing governments (0.319 = 0.175/0.55). In other words, while the average right-wing government permitted an increase in the developable area of the city equivalent to 55% of the initial built-up area, a typical left-wing government only permitted an increase of around 37% (=0.55–0.175).

Although this result is of quantitative importance, we cannot be sure of its meaning, since there may well be many influences on urban growth that we are unable to measure but which are potentially correlated with the partisan identity of the government. Note for instance that, although the equation does identify some of the drivers of growth (i.e., more land is put on the market when there is a shortage at the beginning of the period and where there is plenty of open land, in urban areas, suburbs and on the coast), the explanatory capacity of the model stands at around just 15%.[19]

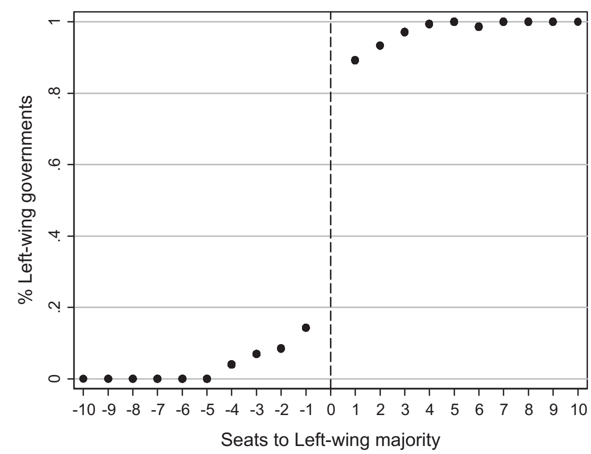

Fig. 1. *%Left-wing governments vs. Seats to left-wing majority*. *Notes*: (1) *% Left-wing government* = proportion of local governments with a left-wing mayor. *Seats to left-wing majority* = number of seats needed for the left-wing bloc to win (if −) or lose a majority of seats (if +).

5.2. Regression discontinuity

Exploring the discontinuity. In order to verify the robustness of these results we employ a more demanding identification strategy, comparing left- and right-wing governments involved in close elections. As explained in the previous section, this is not an easy task in the Spanish case, given the system of proportional representation used and the existence of many coalition governments. To overcome these difficulties we started by looking at close elections in terms of the number of seats won. For this exercise to be relevant, having one more seat should be essential for the partisan identity of the government. Fig. 1 plots the percentage of left-wing governments against the distance in terms of seats between the left- and right-wing blocs: negative numbers indicate the number of seats that the left-wing bloc would need to obtain so as to gain a majority of seats (i.e., to have one more seat than the right-wing bloc), while positive numbers indicate the number of seats the left-wing bloc would have to lose in order to relinquish this majority. Note that the proportion of left-wing governments jumps considerably between −1 and +1 (i.e., after the left-wing bloc wins a majority of seats). The probability of having a left-wing government jumps by approximately 70% at that threshold. This probability also increases when gaining other seats, but the jump in these other cases is much smaller. This suggests that a close-race RD design can be applied in our case by comparing the municipalities in the vicinity of the 50% seat threshold.

However, the fact that under the d'Hondt rule seats are won after only a discrete change in the number of votes means that some of the municipalities in the −1 seats group might be much closer than others – in terms of the number of votes – to gaining the additional seat required to secure a majority (and also that some of the municipalities in the +1 groups are closer than others to losing this). We can use that distance (*% Votes to left-wing seat*

[19] None of the additional controls proved, individually, to be statistically significant at conventional levels, although some did present the expected signs and *t*-statistics above one (e.g., growth seems to be lower in places with a large proportion of homeowners and commuters and higher in places with high rates of unemployment). However, the explanatory capacity of this group of variables is very low, as the *F*-statistics demonstrate.

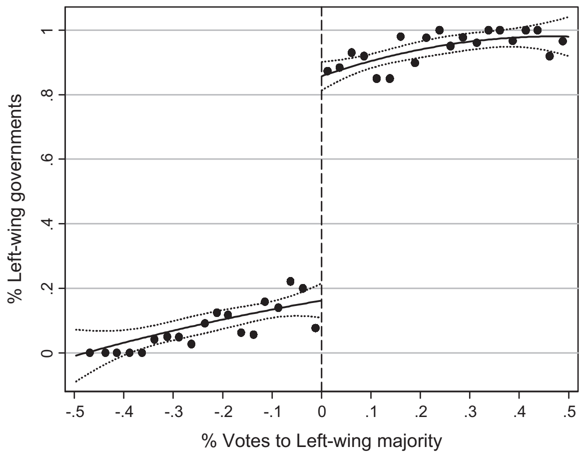

Fig. 2. *%Left-wing governments vs. %Votes to left-wing majority*. *Notes*: (1) % *Votes to left-wing majority* = % of votes that the left-wing bloc should lose (if +) or win (if −) to obtain one seat less or more than the right-wing bloc. (2) Dots = Bin averages. Bin size = 0.025 (2.5% of the vote), selected using the bin test (Lee and Lemieux (2010)). (3) Black line = 2nd order polynomial. (4) Dotted lines = 95% confidence interval.

Table 3
Discontinuity in the probability of having a left-wing Government.

	Two-sided polynomial				Local regression	
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
d(*Left* > *Right*)	0.793***	0.754***	0.705***	0.694***	0.755***	0.727***
	(0.013)	(0.024)	(0.030)	(0.038)	(0.014)	(0.022)
AIC	970.36	647.39	640.95	649.77	–	–
Pol. Order	0	1	2	3	1	1
Bandwidth	100%	100%	100%	100%	25%	12.5%
Obs.	2112	2112	2112	2112	993	536

Notes: (1) Dependent variable is *dLeft* = 1 if the mayor belongs to a left-wing party and 0 otherwise. (2) Explanatory variables: dummy equal to one if the left-wing bloc has more seats than the right-wing bloc (*d*(*Left* > *Right*)), and two-sided polynomial (or locally weighted regression) in the% *Votes to left-wing majority*. (3) Robust standard errors in parenthesis; (4) AIC: Akaike information criterion.
*** Statistically significant at the 99% level.

majority) to identify a sample of left- and right-wing municipalities that are not only close in terms of seats but also in terms of the number of votes needed to lose or win these marginal seats.

However, before reporting the results obtained when using this approach, we should first show that the proportion of left-wing governments also jumps at the 50% seat threshold when we control for the vote distance variable. This is necessary in order to demonstrate that behind the seat discontinuity there is also a genuine vote discontinuity. Fig. 2 reveals this to be the case. The dots are bin averages of the proportion of left-wing governments. The size of the bin is 2.5% of the vote distance and has been selected using the 'bin test' (see Lee and Lemieux, 2010). The black line is a flexible second-order polynomial, fitted separately on each side of the threshold. It is apparent from the graph that the proportion of left-wing governments increases with the vote for the left bloc and that the jump identified in the probability of having a left-government is of the same magnitude as that reported in Fig. 2. The existence of this discontinuity is formally tested in Table 3. Here we present the results of the test when controlling for a two-sided polynomial (using the whole sample) and also when using a locally weighted regression (with the sample corresponding to a restricted bandwidth of 25% and 12.5% of the vote). Note that in any case, the estimated size of the discontinuity is very similar and statistically significant at the 99% level. The results with the optimal polynomial (that of the second order, as indicated by the AIC) and with a locally weighted regression are similar, identifying a jump of 70–75% around the threshold.

Main results. Table 4 presents the RD estimates of the effect of left-wing governments on urban land growth. Panels (a) and (b) display the results with and without the covariates. The first five columns present the results obtained when using the full sample and a two-sided polynomial. The first four columns present the results of the estimation of the reduced form (Eq. (2)) by OLS when controlling for polynomials of different orders. The optimal polynomial order is two, as indicated by the AIC criterion (see Lee and Lemieux (2010)). The size of the effect changes when moving from a polynomial of order zero and one to a second order polynomial, but very little thereafter. The fifth column displays the results of the 2SLS estimation when using the optimal polynomial. The results change little when adding the covariates. The last two columns report the reduced form estimates when controlling for a locally weighted regression. In this case, the impact is also of a similar magnitude independent of the bandwidth, although the level of precision is lower for the smaller sample.

As regards the results, note first that the size of the effect obtained when estimating the reduced form with either the optimal polynomial or the locally weighted regression is of a similar magnitude, around 0.2. The 2SLS coefficient is higher, around 0.3, closely reflecting the fact that it should be equal to the ratio between the reduced form coefficient and the size of the discontinuity estimated in the first stage (i.e., −0.315 = −0.222/0.704). This effect is much greater than that of the OLS one presented in Table 2. This means that a left-wing government would, on average, develop 65% less land than a right-wing government (0.654 = 0.315/0.481).[20] This effect is even more marked than that recorded previously using OLS.

This effect is displayed graphically in Fig. 3. The graph shows 2.5% bin averages and a flexible polynomial fitted to the whole sample. The size of the discontinuity is apparent from the graph. The graph also shows that the slope of the plot is in general negative, suggesting that governments tend to put more land on the market as they move further from the seat majority threshold. This result is consistent with our previous findings that suggest that both left- and right-wing local governments develop more land as local elections become less competitive (see Solé-Ollé and Viladecans, 2012). However, the result should be interpreted with caution, since with an RD design the shape of the non-linear function fitted at both sides of the threshold does not have a causal interpretation. It might simply be that there are some omitted variables that correlate with the vote margin.[21]

Additional results: population heterogeneity. In the first four columns of Table 5 we present the results when dividing the sample according to two proxies of population heterogeneity.[22] The first is an indicator of social polarization in terms of anti- (or pro-) development preferences. We proxy the size of the anti-development group by summing the respective proportions of homeowners, out-commuters and graduates, and that of the pro-development group by summing the proportions of renters, unemployed, and workers in the construction industry. These two variables are expressed in relation to the sample average (=100) and our indicator of social polarization is the absolute value of the difference between them.

[20] To make this calculation we compared the 2SLS results with the % growth of developable land for a right-wing government located closest to the threshold, which in this case was 0.481.

[21] Note, however, that in our sample close-elections do not seem to differ much from non-close elections in terms of observables, as Table A.1 in the Online Appendix shows.

[22] The RD graphs are not included here for reasons of space, but can be found in the Online Appendix.

Table 4
Regression discontinuity: main results.

	Two-sided polynomial					Local regression (reduced form)	
	Reduced form				2SLS		
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Panel (a): Without controls							
d(Left > Right)	−0.191*	−0.214**	−0.222***	−0.201***	–	−0.204***	−0.210*
	(0.103)	(0.087)	(0.103)	(0.104)		(0.094)	(0.109)
dLeft	–	–	–	–	−0.315**	–	–
					(0.146)		
AIC	7492.93	6870.61	6873.74	6876.73	–	–	–
Panel (b): With controls							
d(Left > Right)	−0.187***	−0.224***	−0.241***	−0.225***	–	−0.254***	−0.230**
	(0.061)	(0.085)	(0.102)	(0.067)		(0.093)	(0.107)
dLeft	–	–	–	–	−0.349***	–	–
					(0.115)		
AIC	7382.42	6772.67	6769.79	6775.33	–	–	–
Pol. Order	0	1	2	3	2	1	1
Bandwidth	100%	100%	100%	100%	100%	25%	12.5%
Obs.	2112	2112	2112	2112	2112	993	536

Notes: (1) Dependent variable: Δu, % increase in developable land over the term. (2) 2SLS: *dLeft* as explanatory variable and *d*(*Left* > *Right*) as instrument. (3) Robust standard errors in parenthesis; (4) AIC: Akaike information criterion.
*** Statistically significant at the 99% level.

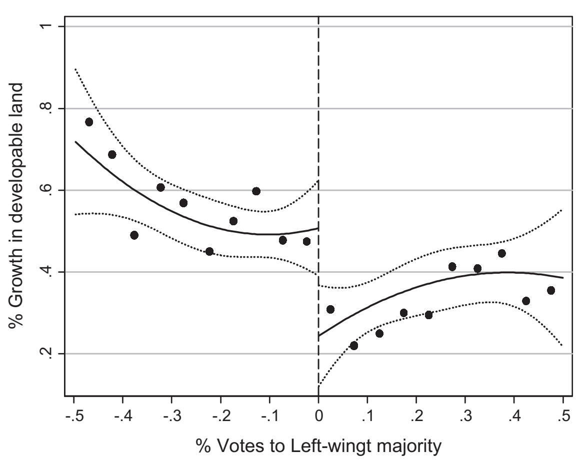

Fig. 3. *%Growth in developable land vs. %Votes to left-wing majority. Notes*: (1) Dots = Bin averages; Bin size = 0.025 (2.5% of the vote), selected using the bin test (Lee and Lemieux (2010)). (2) Black line = 2nd order polynomial. (3) Dotted lines = 95% confidence interval.

The higher the index the more dominant is one of the groups (either the anti- or the pro- growth one) and the more homogeneous is the population. Our expectation (recall the discussion in Section 2) is that the more homogeneous the population the more credible will be the promises the parties make to the median voter, fostering the convergence of policies enacted by right- and left-wing parties. Then, we repeat the RD analysis for the sub-samples of municipalities with social polarization indexes higher and lower than the median.

The results are displayed in the first two columns of Table 5 and suggest that partisan policy differences are much greater in more polarized places, left-wing governments allowing 85% less land to be developed than right-wing parties (recall that this figure stood at 60% for the whole sample). In less polarized communities, the figure is around 35%, but it should be noted that the coefficient is not statistically significant.

The second proxy of social heterogeneity is a measure of the fragmentation of the population between municipalities belonging to the same local area. For each local area we have computed a Hirschman–Herfindhal index of municipal population concentration.[23] A low index value is indicative of a high level of fragmentation, meaning that (for a given population size) the pool of municipalities from which to choose is larger. Our argument here is that fragmentation enhances residential choice facilitating the clustering of population groups with similar tastes, some of which create more homogeneous communities that in turn facilitate policy convergence. Thus, in this case, we expect that the greater the area's fragmentation, the smaller the differences will be between the policies enacted by right- and left-wing parties. The results obtained when dividing the sample between municipalities with values above and below the median value of this index are presented in the third and fourth columns of Table 5. We find that partisan differences are restricted to local areas displaying a low level of fragmentation. In this case, left-wing parties allow 81% less land to be developed than the amount developed by right-wing parties. The differences are much lower (around 30%) in the most fragmented areas but, again, the coefficients are not statistically significant. The results of this analysis suggest that partisan differences in the drawing up of local land use policies tend to occur mainly in the more heterogeneous communities. This finding is similar to that reported by Ferreira and Gyourko (2009) who conclude that (in the case of fiscal policy) there are no partisan differences in suburban US communities.

Additional results: housing demand. As was argued in Section 2, in the case of land use policies, we expect partisan differences to occur only when there is some controversy regarding the desirability of allowing or preventing additional development. Clearly, this only occurs when a municipality experiences a substantial housing demand shock. To verify this intuition we divided our sample in two according to the housing construction growth rate experienced by the local area (here again we draw on the 109 AUDES urban areas plus the 50 provinces) during the previous term-of-office. It is our contention that if the area has grown considerably in the near past, local governments may well forecast that it is likely to grow in the future and, thus, start contemplating the expansion of the amount of developable land to accommodate their forecasts. Our results are presented in the last two columns of Table 5. Indeed, we find that the differences between left- and right-wing governments are more pronounced in rapidly growing

[23] In this case, the definition of local area is the Local Labor Market (LLM), defined using commuting patterns. According to Boix and Galletto (2006), there are 802 LLMs. We computed the Hirschman-Herfindahl index for each of them. We did not use the 109 AUDES urban areas because they do not cover the whole of Spain.

Table 5
Regression discontinuity: population heterogeneity and housing demand.

	(i)	(ii)	(iii)	(iv)	(v)	(vi)
	(a) Social polarization		(b) Area fragmentation		(c) Housing demand	
	Low	High	Low	High	Low	High
RD-OLS	−0.062 (0.057)	−0.498*** (0.122)	−0.951*** (0.241)	−0.174 (0.187)	−0.088* (0.047)	−0.362*** (0.109)
RD-2SLS	−0.088 (0.076)	−0.681*** (0.104)	−1.219*** (0.358)	−0.223 (0.259)	−0.132* (0.075)	−0.514*** (0.155)
% Decrease	35.22	85.86	81.27	36.17	38.01	83.25

Notes: (1) *Demand shock*: % growth in housing construction in the area during the previous four years. (2) *Social fragmentation*: absolute value of the difference between per capita indexes (sample average = 100) of anti-growth population groups (Homeowners + Out-commuters + Graduates) and pro-growth groups (Renters + Aged25–40 + Unemployed + Construction workers). (3) *Area fragmentation*: normalized Hirschman–Herfindhal index of population concentration across municipalities of the urban area. (5) See Table 4.

areas. A left-wing government in one such area will allow 83% more development than a right-wing government located in a similar municipality. In slow-growing areas this number falls to around 37%. The coefficient for this group is statistically significant at the 90% level.

Obviously, the doubt remains as to whether this result is simply a reflection of the previous one regarding the effects of population heterogeneity. We believe this not to be the case as the correlation between the heterogeneity of the housing demand dummies is quite low (less than 5%, in absolute terms) and not statistically significant at any reasonable level.[24] Yet, we must admit that the results obtained when replicating the RD analysis across subsamples should not be extended far. Even if the dummies used to divide the sample in different ways do not appear to be correlated, a correlation might exist with any other variable having an effect on the differences in the behavior of right- and left-wing parties. Thus, the results of the heterogeneity analysis may reflect the explanation we have invoked or many other causes. However, the fact that the three analyses performed point in the expected direction is encouraging.

Additional results: housing market outcomes. Up to this juncture we have shown that the identity of the party controlling the local government does matter as regards the restrictiveness of land use regulations. Left-wing governments – we have shown – allow less land to be converted from rural to urban uses. This begs the questions as to whether this reduced supply of urban land has an impact on housing market outcomes, such as housing construction and housing price growth. To shed some light on this we have data on the amount of developed land for the same period and sample as before, and data on the growth rate of housing prices for municipalities with more than 25,000 residents (see Table 1 for further details). The smaller sample size when using price data is obviously a concern, for several reasons. First, the reliability of the RDD could be affected by the existence of a lower number of observations around the threshold. Unfortunately, little can be done to rectify this shortcoming, since the Spanish government does not disclose information on housing prices for smaller municipalities. However, it should be stressed that the performance of the RDD when using this smaller sample is similar to that when using the larger one. Second, the fact that the analysis is restricted to municipalities with more than 25,000 residents makes it difficult to generalize these results to the smaller municipalities.[25]

In any case, however, identifying the effect of partisan control on housing market outcomes is a difficult task, for at least two reasons. First, and above all in the case of housing prices, the effect of partisan control might extend beyond its impact on land supply. Left-wing governments, for example, might also adopt different fiscal policies, which might have an impact on the demand for locating in the municipality. There is unfortunately no remedy for this and so the results on housing prices have to be interpreted with great care. Second, land supply decisions have a delayed impact on actual development. This means that while it is natural to assume that a given government can modify land use policies during its term-of-office (e.g., convert more land to urban uses), it cannot be so readily assumed that the effects of its policy on construction and prices will make themselves manifest during this short period of time. To overcome this problem, we use a dynamic version of the RDD (see Ferreira et al., 2010). The specification relates the outcome variable of interest in term-of-office t to the discontinuity in both t and $t-1$. In the case of growth in developable land, we have:

$$\begin{aligned}\Delta u_{i,t} = {} & \lambda_t d(\textit{Left seats} > \textit{Right seats})_{i,t} \\ & + \lambda_{t-1} d(\textit{Left seats} > \textit{Right seats})_{i,t-1} \\ & + +f(\%\ \textit{Votes to left-wing majority})_{i,t} \\ & + f'(\%\ \textit{Votes to left-wing majority})_{i,t-1} + \upsilon_{i,t}\end{aligned} \quad (5)$$

As argued above, a given government should be able to implement fully its land use policies during its term-of-office, so we expect that in the case of developable land λ_t will be negative and λ_{t+1} will be zero. In the cases of developed land and housing prices the prediction is not so clear and it could well be that $\lambda_{t+1} < 0$ or even that $\lambda_t = 0$.

The results obtained when estimating this equation for the three variables of interest are presented in Table 6.[26] Columns (i), (iii) and (v) show the static analyses, while columns (ii), (iv) and (vi) present the dynamic analyses. The results for developable land are presented in columns (i) and (ii) and show that the effect is exclusively contemporaneous. The results in columns (iii) and (iv) show that partisan effects are statistically significant at the 90% level in the static model (column (iii)), while lagged partisan effects are significant in the dynamic model (iv), although the size of the coefficient is similar to that of the contemporaneous effect (which is not statistically significant). The effect of parties on housing prices presents a similar profile: housing prices grow higher in areas under left-wing parties, and the lagged effect is stronger (and the only one that is statistically significant). In both cases, the effects are substantial: having a left-wing party in control of local government is

[24] The two heterogeneity variables are negatively correlated; thus, there is greater polarization where there is less fragmentation. The correlation coefficient is around -0.05, although it is not statistically significant.

[25] However, the results can always be compared to the effect of parties on developable land in the same sample (municipalities with more than 25,000 residents). In results not shown here for reasons of space, we found that the main results also hold for these larger municipalities, and that the estimate of the % reduction in the amount of developable land due to a shift from a right- to a left-wing government is of a similar magnitude.

[26] The RD graphs are not presented here for reasons of space, but are included in the Online Appendix.

Table 6
Regression discontinuity: additional outcomes and dynamic effects.

	(i)	(ii)	(iii)	(iv)	(v)	(vi)
	(a) Growth in developable land		(b) Growth in developed land		(c) Growth in housing prices	
	Static	Dynamic	Static	Dynamic	Static	Dynamic
dLeft (*t*)	−0.315	−0.278	−0.071	−0.027	0.026	0.024
	(0.146)***	(0.139)**	(0.047)*	(0.017)	(0.017)	(0.013)
dLeft (*t* − 1)	–	−0.011	–	−0.035	–	0.033
		(0.364)		(0.015)**		(0.014)**
%Decrease (*t*)	65.40	57.72	40.01	15.50	32.08	22.02
%Decrease (*t* − 1)	–	2.28	–	20.10	–	30.28
Pol. Order	2	2	3	3	4	4
Obs.	2112	2112	2112	2112	252	252

Notes: (1) See Tables 3 and 4. (2) 2SLS estimates. (3) All equations have been estimated with the 100% bandwidth and the same controls as before. (4) Columns (v) and (vi) estimated for the sample of municipalities with more than 25,000 residents. (5) *%Decrease* (*t*) and *%Decrease* (*t* + 1) refer to the difference due to contemporaneous and lagged partisan control.

associated with 15% less development in the current term and 20% less development in the future, and with 22% higher growth in housing prices in the current term and 30% higher growth in the future.

Validity and robustness. The validity of the RD design depends on certain assumptions. Firstly, agents should not be able to manipulate the forcing variable. There has been some concern in the literature about this possibility (see Caughey and Sekhon, 2011). Several factors might be behind this result: electoral fraud, differences in the capacity to mobilize resources during a closely contested campaign and, in our case, differences in the capacity to broker coalition deals, either before or after the elections. The first factor can be completely dismissed in Spain, since there are no grounds whatsoever for concern about the possibility of electoral fraud in local elections. The second is equally implausible given the low amount of resources required to run a local campaign. Moreover, as Caughey and Sekhon (2011) note, the manipulation of the forcing variable is more feasible in a two-party system with very sophisticated polling systems, where the level of uncertainty regarding the election is greatly reduced. However, they claim that manipulation is less feasible in proportional electoral systems and in places where campaigning is not especially sophisticated. This description matches Spain's local elections perfectly. As for the last factor, we should stress that pre-electoral coalitions are extremely rare in Spain, as a result of the incentives generated by a system based on proportional representation. Post-electoral coalitions do constitute a potential threat to our empirical strategy, but to avoid it we have worked with ideologically linked blocs of parties rather than with actual coalitions. In any case, we have performed several checks to discard the possibility of manipulation (see the Online Appendix). The histogram of the vote distance and a more formal test (McCrary, 2008) show that the density of the forcing variable is continuous at the threshold. We also report discontinuity tests showing that none of the pre-determined covariates is affected by the discontinuity.

We performed a number of additional analyses in order to demonstrate that our findings are not influenced by any particular methodological decision (see also the Online Appendix). Firsly, we repeated the analysis but this time we eliminated from the sample those municipalities with at least one seat allocated to a local party. The results were virtually unchanged. Secondly, we undertook the analysis using only those municipalities in which the two main parties obtained most of the votes, i.e., a situation that resembled a bipartisan system. Here, the discontinuity was greater than before, but the estimated effect was very similar. Thirdly, we restricted the sample to include just coalition governments, so as to show that the discontinuity is not an artifact created by the fact that our sample contains more majority governments than coalitions. Here the jump was around 50%, which is lower than the 70% reported for the whole sample. However, the treatment effect is of the same magnitude. Fourthly, we repeated the analysis using an alternative measure for the voting distance needed to win or lose a majority of seats. So far the distance used has been computed on the assumption that the votes won/lost come/go from/to abstentions. Now we adopt a measure that assumes that these votes might come/go not only from abstentions but also from the other ideological bloc. The results were again unchanged.

Finally, we also present results when comparing subsets of left- and right-wing governments. So far in the discussion we have implicitly considered all parties in one ideological bloc as being equivalent. However, the discussion in Section 3 suggests that some left-wing parties are more anti-growth than others (e.g., IU, the former communists, closely linked in Spain with the environmental movement), and also that some right-wing parties are more pro-growth (e.g., PP, closely linked with the complete deregulation of the land market). The results show that there are no differences between Center–Left (CL) and Center–Right (CR) parties, but that the differences between Center–Left (CL) parties (the main left-wing party, PSOE, in most instances) and Right–Right (RR) parties (the right-wing party, PP in most instances) are larger than those presented before (see Table A.4 in the Online Appendix).

6. Conclusion

This paper has analyzed whether the ideology of the party controlling the local government has an influence on a municipality's land use policies. In so doing, we have drawn on a new database containing information about the amount of land converted from a rural to urban use by Spanish municipalities in the period 2003–2007. To identify the effect of the country's political parties we have used a close-election regression discontinuity design, amended to account for the specific institutional traits of Spain's local political system. Our method has involved the comparison of governments controlled by left-wing and right-wing parties that are close to holding a one seat majority in the council, while controlling for a function of the distance in terms of number of votes to losing or winning a majority. Our results suggest that left-wing governments have a considerable influence on land use policies. Left-wing governments that are close to winning-losing power allow 65% less land to be developed than comparable right-wing governments. The effects of left-wing parties are particularly pronounced in more heterogeneous communities and in places facing greater housing construction growth rates. It would seem to be the case that it was in these places that the conflict between pro- and anti-growth groups was most pronounced during this period and the consensus regarding the desirability of urban development most difficult to achieve.

In addition, we have shown that municipalities controlled by left-wing parties also build less and present higher housing price growth rates. These effects, which are substantial, persist moreover into future terms of office. We recognize that our housing price results need to be considered with a degree of caution given the size and characteristics of the sample. Furthermore, we should point out that left-wing governments might affect housing prices via other channels. Yet, it is our belief that, when considered together with the results on developed land, the effect on housing prices could well be attributable to overly restrictive land policies. For instance, it should be noted that the effects of left-wing parties on both developed land and prices are similar in magnitude. Moreover, the impact on prices is positive, which rules out the possibility of the supply restriction effect being undone by a demand effect (e.g., left-wing fiscal policies that appeal to low-income households). Whatever the case, further research is needed to assess whether the social benefits provided by land use constraints (or the careful design of social housing) are able to undo or compensate for the housing affordability problems generated by these price increases.

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Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at http://dx.doi.org/10.1016/j.jue.2013.07.003.

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