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Patterns of Land Use, Zoning and Economies of Scale in Cities: Three Essays on Urban Economics for Brazil

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Tese apresentada ao Programa de Pós-Graduação em Economia (PIMES) da Universidade Federal de Pernambuco, como requisito parcial para obtenção do título de Doutor em Economia.

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RESUMO

Essa tese está dividida em três ensaios empíricos no campo de economia urbana, analisando o caso das cidades brasileiras. O primeiro ensaio utiliza uma base de dados inédita contendo informações sobre 98.198 lotes individuais da cidade do Recife para investigar os determinantes dos diferentes usos do solo urbano. Através de uma abordagem espacial não paramétrica, observa-se que existem consideráveis efeitos espaciais locais. Além disso, notamos algumas regularidades na distribuição espacial do uso do solo urbano. As infraestruturas de transportes tendem a atrair o desenvolvimento de atividades comerciais, enquanto as amenidades naturais incentivam o uso residencial. Já o segundo ensaio, concentra-se em avaliar o impacto das políticas de zoneamento urbano sobre o mercado habitacional. Utilizando o método de pareamento e um conjunto de dados que abrangem todos os municípios brasileiros, evidencia-se que as leis de zoneamento provocam um acréscimo que varia entre 5.4% a 6.3% nos aluguéis médios, mas não afetam o crescimento no número de casas. Adicionalmente, conduzimos uma série de análises de sensibilidade que mostram que esse resultado não é simplesmente guiado por variáveis não observáveis. Por último, o terceiro ensaio investiga o impacto da secessão de municípios sobre os gastos públicos locais através de uma abordagem de Diferenças em Diferenças. Os resultados indicam que os municípios que experimentaram um processo de secessão aumentaram as suas despesas per capita com capital em cerca de 14.7%. Além disso, mostramos evidências que sugerem que esse aumento pode ser explicado por perda de economias de escala e pelo aumento do comportamento *rent-seeking*.

Palavras-chaves: Análise do Uso do Solo Urbano. Restrições de Uso do Solo. Avaliação de Política Pública. Secessão de Munícipios. Finanças Regionais e Urbanas.

ABSTRACT

This thesis is divided into three empirical essays in the field of urban economics, analyzing the case of Brazilian cities. The first essay uses a novelty microdata of 98,198 individual parcels in the city of Recife to investigate the determinants of the urban land use. Through a semi-parametric spatial econometric approach, it is observed that there are considerable local spatial effects. In addition, we note some regularities in the spatial distribution of urban land use. Transport infrastructures tend to attract the development of commercial activities, while natural amenities encourage residential land use. The objective of the second essay is to evaluating the impact of zoning ordinances on the housing market. Using matching methods and a dataset covering all Brazilian municipalities, we evidenced that zoning generates an increment ranging from 5.4% to 6.3% on the average rents, but do not affect the house growth. In addition, we conducted sensitivity analyses that suggests that this result is not simply driven by unobservable confounders. Finally, the third essay investigates the impact of municipal secession on local public expenditures through a Differences-in-Differences approach. The results indicate that those municipalities that experienced a secession process increased their per capita capital expenditures by 14.7%. In addition, we show evidence that suggests that this increase in expenses can be explained by a reduction in economies of scale and rent-seeking behavior.

Keywords: Urban land Use Analysis. Land Use Restrictions. Public Police Evaluation. Municipal Secession. Urban and Regional Finance.

SUMMARY

Overview

This thesis consists of three non-directly related essays in the field of empirical urban economics. It is intended to contribute to the literature by bringing discussions that were never designed for the Brazilian case and have fundamental importance in understanding the dynamics of cities and the well-being of citizens. Specifically, the goal of this study is to understand the patterns of urban land use for a specific Brazilian city, investigate the effects of zoning ordinances on the Brazilian housing market, and, finally, to analyze the impacts of the municipalities secession process.

The first essay investigates the determinants of urban land use for the city of Recife, Brazil. Using a novelty microdata of 98,198 individual parcels and a locally weighted multinomial logit, we show that there is considerable local spatial effects, which indicates that parametric approaches may be inadequate in an intra-urban context. Additionally, some general patterns of land use are identified: transport infrastructures (which include main avenues, roadways and subway stations) tend to attract commercial activities, while natural amenities (including the public open areas, the beach and the river) tend to encourage residential development. Finally, it is evidenced that, on average, the probability of unimproved lots is higher in areas near the city historical center and the verticalization is greater in areas near the city center and to the natural amenities.

Due to recent urbanization and the increased population density around major Brazilian cities, the real estate market in the country suffered a great appreciation. In addition to demand-related factors, mechanisms with potential to limit housing supply may have also affected the price increase in that market. The objective of the second essay is to investigate the impact of zoning orders (land use regulations) on the average rent prices and the real estate market growth in Brazil. Through an intercity analysis and using matching methods, it was observed that zoning generates an increment ranging from 5.4 to 6.3 percent in average rents, but does not affect the house growth. In addition, we conducted sensitivity analysis that suggests that this result is not simply driven by unobservable confounders. Our evidence suggests that even being a usually well-intentioned policy, zoning tends to generate social costs that need to be taken into account in the analysis of the real estate market in Brazil. In the 1990s, there was a growing process of administrative decentralization that culminated in the creation of 1,016 new municipalities in Brazil. The aim of the third chapter is to verify the impact of the municipal secessions on the public expenditures and its association with economies of scale. Based on a Differences-in-Differences methodology, the obtained set of evidence indicates that those municipalities that underwent a secession process increased their per capita capital expenditures by 14.7 percent. In addition, we show evidence that strongly suggests that this increase in expenses can be explained by a reduction in economies of scale and rent-seeking behavior.

Chapter 1 - What Drives the Patterns of Urban Land Use in a Developing Country? The Role of Transport Infrastructures and Natural Amenities

1.1. Introduction

Urban land use plays a key role in urban economics since it reflects the allocative decisions of firms and households (Duranton and Puga, 2015). Firms define their geographic location based on cost minimization, while households decide where to live maximizing their utility and considering the tradeoff between commuting costs and housing consumption. Thus, the spatial distribution of urban land use is a consequence of interactions that occur between individuals, businesses, the environment, and transport systems and the land usually goes to the sector with greater willingness to pay for its use, i.e. the highest bidder (O`Sullivan, 2012). Understanding the patterns of urban land use is fundamental since several urban problems are associated with land use, such as traffic congestion (Sarzynski et al., 2006, Antipova et al. 2011), violence (Browning et al., 2010, Twinam, 2017, Law et al., 2015) and environmental deterioration (Arnott et al., 2008). From a practical point of view, knowing the spatial distribution of land use is useful to explain the geography of social life in the city, support the proper allocation of public services and infrastructure (such as schools, hospitals and roads) and help inform the formulation of more efficient urban policies (such as master plans, zoning laws and housing programs).

In recent decades, the patterns of urban land use of Brazilian cities experienced significant changes as a direct consequence of both the growth of cities and the intense and quick urbanization process. The residential capital stock in Brazil practically doubled in the 1990-2008 period, an increase of 98.3%, well above the real GDP growth in the period of 80.5% (IPEA, 2008). Urban space grew horizontally and vertically, although the second type of expansion occurred with a greater intensity. For example, in 2000, about 56.5% of the population living in metropolitan regions resided in the central city, but, in 2010, this proportion fell to 54.8%, indicating a further growth of suburban cities. In relation to verticalization, there was an increase of 11.92% in the proportion of dwellings classified as apartment buildings from 2000 to 2010, contrasting with the 2.68% decrease in the proportion of dwellings classified as traditional houses (Brazilian Demographic Census, 2010). It is necessary to recognize that these phenomena are also observed in other Latin American cities and are related to the rise of gated communities (Borsdorf et al. 2007) and the decentralization of employment and formation of polycentric structures (Fernández-Maldonado et al. 2013). In addition to the recent changes in the spatial configuration of cities, another important characteristic of developing countries is the huge informal housing market. Although the proportion of individuals living in the informal market has declined over time, it is estimated that 21% of the urban population of the Caribbean and Latin America still live in informal settlements, which represents a population of approximately 105 million people. (UN-HABITAT, 2016). This strong informality may be a consequence of several factors such as intense rural-urban migration, scarcity of housing supply, income inequality and rigid land use regulations (Cavalcanti et al. 2017).

Considering the importance of the subject and the recent modifications in the urban configuration of Brazilian cities, the objective of this paper is to investigate the determinants of urban land use analyzing the case of Recife, an important Brazilian urban center and the oldest capital of the country. More specifically, using a novelty microdata set containing information on 98,198 individual lots and employing a semiparametric spatial discrete choice model, we will associate the categories of land use (residential, commercial, multifamily residential, mixed use and unimproved) with different kinds of urban attributes: accessibility of the lot, availability of natural amenities and, finally, local transport facilities. Our study area is particularly interesting for two reasons: first, the city does not have a zoning ordinance that restricts the allocation of land for different categories of use and, second, Recife has natural amenities that generate strong heterogeneities in different spaces. Empirically, these singular characteristics allow us to investigate whether a competitive market allocation of land leads to a clear segregation of uses in the urban space and how natural amenities can shape the city's spatial configuration.

There is a significant body of empirical studies related to ours that can be subdivided into three groups. First, there arepapers that analyze how the conversion of forest or agricultural land into urban use works (Carrión-Flores et al., 2009; Li et al., 2013; Chakir and Parent, 2009; Chakir and Le Gallo, 2013). Second, studies that investigate the changes of land use in an intra-urban context, attempting to understand the urban attributes that affect land conversions for residential, commercial or industrial use (Verburg et al., 2004; Páez, 2006; Braimoh and Onishi, 2007; Wang et al., 2011; Wrenn and Sam, 2014; Bhat et al., 2015). Finally, the most similar ones to ours, that seek to determine the factors that affect the allocation of land to different urban uses, like commercial, residential, industrial and mixed use (McMillen and McDonald, 1999; McMillen and Soppelsa, 2015; Jacob and McMillen, 2015). This last type of research is scarcer, since most cities have a zoning ordinance that tend to interfere and alter the allocation of land far from what would be determined by a competitive market, making it difficult to investigate the natural patterns of land use (McDonald, 2006).

This literature highlights the role of accessibility - commonly measured by the proximity to the Central Business District (CBD) - and the role of transport infrastructures in explaining the patterns of urban land use. For example, Páez (2006) found that the implementation of Union City BART station (California) changes the land use in areas close to the new facility. Wang et al. (2011), analyzing the city of Austin (Texas), evidenced that proximity to CBDs encourages the development of office buildings but discourage commercial and industrial development. Additionally, they showed that areas close to major arterial roads are more likely to host commercial and industrial conversion. Using a detailed data set of 474,190 parcels from Cook County (Illinois), McMillen and Soppelsa (2015) show that residential land use tends to be smaller in regions close to the CBD, to train lines and to major streets. The availability of natural amenities and the biophysical quality of the soil are also recognized as important variables to explain the urban land use (Verburg et al., 2004, Wang et al., 2011).

A concern that is common to these empirical studies is the development of discrete choice models incorporating the spatial dependence hypothesis. It is well recognized that patterns of urban land use are strongly dependent on the neighborhood, i.e., it is undoubtedly a spatial process. Commerce, industry and other business activities tend to develop together to take advantage of agglomeration economies, either through input-output linkages, knowledge spillovers, consumption externalities, or labor pooling (Billings and Johnson 2016). Residential areas also have economies of agglomeration, mainly because they can share the same public infrastructure, like sewage, telecommunications network and electricity. Considering this, the most recent studies incorporated spatial dependence into discrete choice models in two different ways: through parametric approaches which estimate a global measure of spatial dependence, in line with Anselin (1988) and through the application of a semiparametric approach based on extensions of the Locally Weighted Regression (LWR) and allowing spatial non-stationary relations among the variables. In the first segment, Chakir and Le Gallo (2013), Wang et al. (2014) and Bhat et al. (2015) found that land use change has a positive autocorrelation, indicating that conversions to a specific use affect the conversions of neighboring parcels for the same use. As for semiparametric approaches, Wang et al. (2011), Wrenn and Sam (2014) and McMillen and Soppelsa (2015) show that, although there is a huge spatial variation in the magnitude of factors that affect the allocation of lots to different uses within a particular city, parcels that are geographically close tend to develop very similarly.

At this point, it should be noted that most empirical studies are conducted for North American cities and, therefore, very little is known about the determinants of urban land use in the developing world. The existing evidence indicates that the patterns of land use are sensitive to distinct institutional contexts, perhaps because of the different degrees of urbanization between countries. For example, for Lagos (Nigeria), Braimoh and Onishi (2007) found that residential development tends to be higher in areas close to the Central Business District (CBD), while McMillen and Soppelsa (2015) show just the opposite. In addition, much of the previous literature has prioritized the theoretical development of spatial discrete choice models, neglecting the empirical investigation of land use determinants. This created a large gap between methodological advances and empirical land use analyses. Therefore, given the recent changes in the spatial configuration of cities in developing countries - mainly the increased urban sprawl, the verticalization and the huge informal housing market - and the lack of empirical studies analyzing urban land use in these countries, it is extremely important to develop new research in the area.

We intend to contribute in this sense, analyzing the determinants of land use for the case of Recife, a large Brazilian metropolis. Since Recife has urban trends that are similar to other Latin American cities' and the spatial distribution of land use is market-oriented, we believe that, with some attention, the results can be expanded to other cities of the developing world. Our paper also innovates by analyzing the factors that affect land allocation for tall buildings and for unimproved lots, contributing to the literature on the determinants of verticalization (Barr, 2010, Ahlfeldt and McMillen, 2017) and to studies that analyze vacancy rates (Morandé et al., 2010; Nadalin and Igliori, 2016). Despite the strong increase in verticalization in Brazil, the empirical evidence about its determinants is practically nonexistent.

The rest of the paper is organized as follows: in section 2, we discuss the main theoretical predictions regarding urban land use, in section 3 we briefly present the urban trends and characteristics of Recife and describe the database, section 4 details the empirical strategy, section 5 presents the results and the discussion, section 6 presents some aspects related to the predictive capacity of the empirical models and, finally, section 7 brings the final remarks.

1.2. Theoretical Background: the landowner decision and the bid-rent functions

The patterns of urban land use are determined by the economic interactions between the landowner and the different sectors of the urban economy (housing, manufacturing, commerce, offices and others). From the landowner perspective, they are confronted with the decision to allocate parcels of land - with constant quality and size - for different categories of use. In addition, it is assumed that the landowner is risk-neutral and has the objective of maximize the present discounted value of all rental flows of land parcels. Considering these assumptions, the decision rule that stems from a dynamic optimization problem (described in detail, for example, by Lubowski(2002)) is to choose, in each time period t, the land use that generates the highest net return. The net return on land in use k is equal to the return generated by the parcel of land allocated in use k minus the potential opportunity cost of conversion. k is an element of the total set of different land uses, denoted by $K = (1, \ldots, k)$. A landowner with land in use j will choose the use k at time t that satisfies:

$$ArgMax_{k}\left(R_{kt} - rC_{jkt}\right) \geq R_{jt} \tag{1}$$

Where R_{kt} represents the expected return of the land in use k at time t, r is the discount rate and C_{jkt} is the expected marginal cost of converting land in use j to use k at time t. Therefore, the landowner's optimal decision is to convert the land in use j to the use k if the return of the second category exceeds the return of the first category (after deducting the conversion costs), not convert if the return in use j exceeds the return in use k, and finally, the landowner can be indifferent between the different categories of land use. The description of the landowner's behavior is useful to illustrate one of the main results of urban economics: the land goes to the urban sector with greater willingness to pay for its use, or, in other words, the highest bidder (Fujita, 1989; O`Sullivan, 2012). Usually, the amount the sectors are willing to pay for land is called bid rent.

The bid-rent is a function of several factors, ranging from the biophysical quality of the soil, to the availability of transport infrastructure and other public goods. Considering the close link between bid-rent and land use, in the present paper we will associate the categories of land use with three different types of urban attributes: the degree of accessibility of the lot (measured as the distance to the CBD), the proximity to transport facilities and, finally, the availability of natural amenities.

About the first aspect, the traditional Alonso-Muth-Mills model (Alonso, 1964; Muth, 1969; Mills, 1972) is useful to predict how land will be allocated across different regions of the same urban area. One of the main conclusions of the model is to show that, in equilibrium, the price of land decreases according to the distance from the CBD (Brueckner, 1987). As a consequence, when deciding where to live, households face a trade-off between consuming more housing (living farther from the CBD) or getting closer to the workplace and reduce the commuting costs (living closer to the CBD). That is, residential development in areas close to the CBD is encouraged by accessibility and, on the other hand, discouraged by higher land prices. The willingness to sacrifice accessibility for space is heterogeneous among different households and varies according to income (Duranton and Puga, 2015). In addition, because land is more expensive in areas close to the CBD, housebuilders choose to replace land with more capital and thus produce taller buildings (Fujita and Thisse, 2013; Ahlfeldt and McMillen, 2017). With regard to business - such as commerce, industrial activities and offices – these tend to benefit from the proximity to the CBD, due to potential agglomeration gains.

Transport infrastructures also play a key role in determining the patterns of urban land use. As described by McDonald (2006), a new transportation project changes the bid-rent of agents to match the savings in transport costs, and, consequently, affect land allocation to different uses. First, the business sector strongly benefits from the proximity to transport infrastructures: it facilitates the flow of goods, reduces freight costs, and improves access to the consumer market (O`Sullivan, 2012). Thus, the bid-rent for commercial or industrial uses is expected to be greater in places closer to the transport facilities, since these areas tend to be more profitable. In a recent paper, Billings and Johnson (2016) show that proximity to transport infrastructures explain a great part of the industrial agglomeration in the Denver-Boulder-Greeley CMSA. With respect to the effects on residential land use, transport proximity can either increase or reduce the bid-rent of households. Easier access to transport infrastructures has the benefit of reducing costs related to commuting, improving access to amenities and public services and reducing shop trips distance (McDonald, 2006). However, in areas very near to infrastructures, a reduction in the bid-rent may occur due to the predominance of negative externalities such as congestion, pollution and noise. The bid-rent function for mixed land use behaves in a very particular way. Mixed land use is a response to situations where commuting costs are relatively high and agglomeration forces are low (Wheaton, 2004; Fujita and Thisse, 2013). Under these circumstances, firms tend to locate more dispersely and households are encouraged to work near their residences, encouraging mixed use conversions. Thus, it is expected that mixed use be associated with locations farther from transport facilities and with the peripheral area of the city (in relation to CBD), where the capacity of agglomeration gains is limited.

It is also necessary to consider the role of environmental quality in household and business location decisions. In this sense, Cho (2001) and Brueckner et al. (1999) expanded the Alonso-Muth-Mills model by incorporating the availability of natural amenities into the utility function of households. In the residential sector, is expected that bid-rent will be greater in places closer to natural amenities. In such places, families benefit from having access to more beautiful views, a cleaner air or more leisure and recreation options. However, except for specific sectors such as touristic and gastronomic, firms do not benefit directly from being closer to natural amenities and, therefore, do not tend to have a greater bid-rent in areas with good environmental quality.

In summary, the theoretical results point that the degree of lot centrality and the proximity to transport infrastructures tend to increase the bid-rent of the business sector and generate an inconclusive effect on household bid-rent (which will depend on the preferences of each of them). Areas close to the CBD also tend to have taller buildings and a smaller proportion of mixed land use. On the other hand, the household's willingness to pay for land use tends to increase in the vicinity of natural amenities, while the bid-rent of the business sector is only affected by this attribute in very particular cases.

1.3. Context and Data

1.3.1. Urban characteristics of the study area

Our study area comprises the municipality of Recife, the main city of the Metropolitan Region of Recife (RMR) and the capital of the Pernambuco state. The RMR is one of the main urban agglomerations in Brazil, being the richest metropolis of the North and Northeast regions (with a GDP of R$ 75.8 billion, according to Brazilian Institute of Geography and Statistics, 2015) and the eighth richest in the country. The metropolitan region has a population of about 3.94 million of residents and Recife concentrates 41.2% of this contingent.

In recent decades, the city has undergone strong modifications in its urban structure, and, like other Latin American cities, it experienced processes of intense verticalization, urban sprawl and economic decentralization in relation to its historic center. As shown by Silveira-Neto (2016), in the 1991-2000 period, the districts closer to the center of Recife grew by an average of 5.5% while the peripheral districts grew by 19.9%. This tendency of sprawl was attenuated in the most recent period (2000 to 2010), where both areas grew by about 8%. However, despite being an old city (founded in 1537) and having experienced some urban sprawl and employment decentralization, Recife remains with an eminently monocentric structure. In a recent paper, Belmiro et al. (2017) identified the centers and sub-centers of Recife and showed that the peaks in employment density are concentrated in a single region that is located in the intermediations of the historic center of the city, indicating the presence of a monocentric pattern. The authors also show that there is a negative and convex relationship between density and distance to the CBD, which reinforces the monocentric structure of Recife and confirm the Alonso-Muth-Mills model predictions.

Although the urban sprawl has been limited, the city grew to the sky: from 2000 to 2010, there was a 14.5% increase in the proportion of dwellings located in apartment buildings (Brazilian Demographic Census, 2010). The more vertical growth of the city is a consequence of two factors: the lack of urban infrastructure in peripheral regions (such as sewage), which inhibits land use in suburban areas and, in second place, the higher income and the massive housing subsidies, which have led to increases in demand for areas with good infrastructure and closer to the city center (Silveira-Neto, 2016). Both factors valorized the land located in the qualified areas, encouraging the exchange of land by capital and, thus, the construction of taller buildings. In addition to these factors, the zoning code may also have facilitated the recent verticalization of Recife.

Unlike major cities such as São Paulo and Brasilia, the land-use legislation of Recife[1] does not impose any type of direct constraint on the allocation of land to particular uses. For example, there are no legal divisions between residential and commercial areas. The zoning code of the city is characterized by restricting only the intensity of use (through maximum floor-area ratios and height limits) and the land-use in areas of environmental protection, of social interest and near of historical and cultural buildings. For this reason, the urban planning legislation is criticized for being too permissive for any kind of harmful and adverse land use (Medina, 1997). In this way, market interactions, rather than a government planner, determine great part of the spatial distribution of land use in the city.

1.3.2. Data

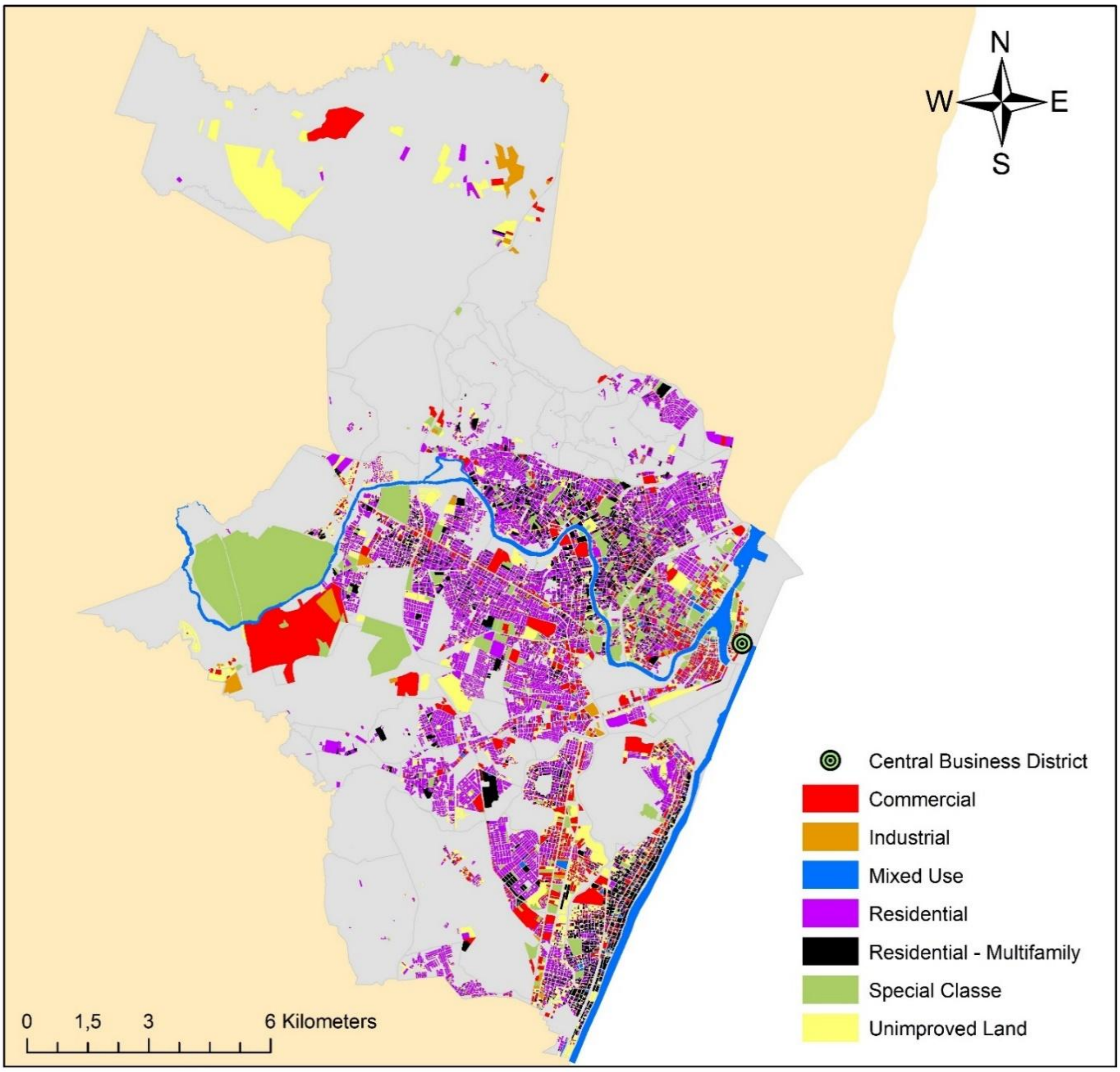
[1] The land-use legislation of Recife comprises the following ordinances: Law 16,176/1996, Law 16,719/2001 and Law 17,511/2008.

Considering the purpose of explaining the factors that are associated with a particular urban land use, we will adopt a novelty and official database that contains information on 98,198 individual lots in the city of Recife. This includes all parcels with property rights, leaving aside the informal ones. The database contains the following variables: area, geographic coordinates, land status (active or unimproved), number of floors, year of building construction and the form of land use. The database was updated in 2013 and is constructed by the local government through the Geographic Information System of Recife (RESIG). Initially, the parcels are divided into the following categories of land use: Commercial[2] (7.78%), Industrial (0.24%), Mixed Use (0.98%), Residential (70.52%), –Mutifamily Residential (6.69%), Unimproved (11.27%) and Special Classes (2.49%). Mixed use refers to lots that cover simultaneously residences and business activities. The last type of land use refers to schools, hospitals, religious temples, public buildings and other unrecorded institutions. For descriptive purposes, Figure 1 shows the spatial distribution of lots by categories of land use.

Although a simple visual inspection does not reveal specific patterns, the following general characteristics can be observed in the Figure 1: the predominance of multifamily residential lots on the seashore, the existence of residential agglomerations in the suburban areas, and the concentration of commercial activities in the vicinity of the CBD. Additionally, it is noteworthy that there are few lots with industrial and mixed use. The first fact is that Recife is a predominantly service-oriented city, with around 85% of GDP coming from this activity (Brazilian Institute of Geography and Statistics, 2015). The industrial plants of the Metropolitan Region are located in other municipalities, excluded from our study area. Low mixed use can be a consequence of both the high agglomeration forces that exist in the city - since Recife remains with a strong monocentric structure - and the lack of an urban planning tradition that values mixed land use.

[2] This category includes a wide range of private business activities: retail stores, offices, parking's, gas stations, financial institutions, hotels and others.

Figure 1 - Spatial distribution of lots by categories of land-use.



Source: Constructed based on Geographic Information System of Recife, 2013.

The explanatory variables were constructed with a geographic information system (GIS) program that enables calculating the distance in relation to each particular lot. Initially, we had taken as CBD the traditional historic center of Recife, known as *Marco Zero* or *Rio Branco* square. According to official records, this place is considered the starting point of Recife. As shown by Belmiro et al. (2017), the city's employment density peak is located in the vicinity of the historic center and, in addition, about 59.75% of all commercial parcels of the city are located within a radius of 5km. In relation to transport variables, we included in our analysis the distance (in meters) in relation to the three major infrastructures of the city:

the subway stations, the main avenues[3] and the roadways. Finally, the natural amenities include the *Capibaribe* River, the beach and the main public open spaces[4]. Table 1 describes the summary statistics of explanatory variables.

Table 1 – Summary Statistics of Lots in Recife

	Mean	**Std. Deviation**	**Maximum**	**Minimum**
Dist. CBD	6,131.2	2,335.3	18,297	26.20
Area (m²)	579.6	17,845.8	3,968,246	0.020
Dist. Roadways	3,090.9	1999.8	7912	14.624
Dist. Main Avenues	506.1	542.7	8,739	1.198
Dist. Subway Station	2,575.3	1,890.7	14,844	14.01
Dist. Public Open Areas	1,039.0	763.2	11,998	1.20
Dist. Beach	4,830.5	2,457.0	18,544	49.925
Dist. *Capibaribe* River	2,465	2,143.0	9,765	6.282

Note: The number of observations is 98,198 and the distances are measured in meters.

1.4. Empirical Strategy: the Locally Weighted Regression

As a way to associate the categories of land use with different urban attributes, we will apply a nonparametric strategy of spatial modeling: the locally weighted regression (LWR). This approach is useful in an intra-urban context to understand nonlinear relationships and to enable spatial heterogeneity across the estimated coefficients. In the study of urban land use, it is very common that the relationships between the variables are different in each particular geographic area. For example, as discussed in subsection 2.2, residential development may be both attracted and driven away by proximity to transportation infrastructures. Locations very close to transport facilities tend to be more crowded, polluted and noisy, what discourages residential land use. However, locations relatively farther (but still close) tend to attract families that wish to benefit from greater accessibility provided by transport facilities. When it is recognized that there is a significantly spatial variation in a particular phenomenon, a common practice in applied econometrics is to estimate different functions for different geographic areas. The LWR simply formalizes this heuristic approach (McMillen and McDonald, 2004). Through the LWR, it is possible to get different coefficients for each

[3] We considered the main avenues those classified as arterial routes by the city's transit agency. In the Appendix, we list these main avenues.

[4] We use only the thirty largest open spaces in the city, which includes parks, squares and green areas. In the Appendix, we list these areas.

particular observation. Thus, unlike the traditional parametric techniques, the method allows spatial non-stationarity. The usual procedure for estimation is the repetition of weighted least squares for each observation. However, in the case of discrete dependent variables, the estimation method involves the maximization of likelihood functions (Tibshirani and Hastie, 1987). In this paper, as we are interested in understanding the determinants of urban land use, the dependent variable will be discrete with M categories. In this way, like McMillen and McDonald (1999) and Wang et al. (2011), we apply a nonparametric multinomial logit model. Considering that there are M categories of land use, the local estimates are obtained by maximizing the following pseudo log-likelihood function for each observation:

$$lnL_i = \sum_{j=1}^{N} w_{ij} \left[I_{0j} \ln\left(P_{0j}\right) + \cdots + I_{Mj} \ln\left(P_{Mj}\right) \right] \tag{2}$$

Where I_{Mj} is an indicator variable that takes value one when the category M is chosen and takes value zero, otherwise. The term P_{Mj} is the probability of choosing the alternative M and is defined as follows (normalizing the base alternative to zero):

$$P_{Mj} = \frac{\exp\,(\beta'_{Mi} x_j)}{1 + \sum_{s=1}^{M} \exp\,(\beta'_{si} x_j)} \tag{3}$$

In equation (3), the term x_j is a vector of k explanatory variables that have potential to affect the urban land use, and β_{Mi} is the associated vector of coefficients. The maximization of equation (2) produces n distinct estimates of the vectors $(\beta_1, \ldots, \beta_k)$ for each category of land use M, which we denote by β_{Mik}.

Before maximizing equation (2), it is necessary to establish a way of associating observations that are spatially close in order to obtain the spatial weights, w_{ij}. This is done through a Kernel weighting function, which, having as inputs the geographic coordinates of i and j, generates the spatial weight w_{ij}. This type of nonparametric process follows the most fundamental law of spatial econometrics: the shorter the distance between i and j, the greater the weight given to these points. A range of kernel functions can be arbitrarily used at this stage. For example, a function that is widely used is the tri-cube:

$$w_{ij} = \left[1 - (d_{ij}/d_{iQ})^3\right]^3 I\left(d_{ij} < d_{iQ}\right) \tag{4}$$

Where d_{ij} is the Euclidean distance between point i and point j (based on geographic coordinates), d_{iQ} is the distance between i and the Q-th nearest neighbor and $I\left(d_{ij} < d_{iQ}\right)$ is an indicator function that equals one only when the condition is satisfied. Thus, this function puts zero weight in the points that have a distance to the target point i higher than the Q-th neighbor and put a non-zero weight that decreases with the distance to those points that are below the threshold. The number Q is called the window size of the kernel function and is conceptually similar to the bandwidth parameter found in other kinds of kernel functions. These parameters determine how fast the weights decrease with the distance (McMillen and McDonald, 2004). Although the choice of the kernel function is not sensible, the choice of the window size or bandwidth is critical to the estimation process (Wheeler and Paéz, 2010). Usually, the choice of these parameters can be made based on an interactive method of Cross-Validation (CV), through the maximum value of the log-likelihood function or based on the Akaike Information Criterion (AIC).

Finally, it is important to note that the estimated vector of β_{Mik} does not correspond to the marginal effects of an increase of an explanatory variable k on the probability of selecting alternative M. In a multinomial logit approach, the marginal effects are given by (Croissant, 2011):

$$\frac{\partial P_{Mi}}{\partial x_{ik}} = P_{Mi}\left(\beta_{Mik} - \sum_{S} P_{Si}\beta_{sik}\right) \tag{5}$$

The second term of the expression in brackets represents the weighted average of the coefficients for all alternatives. The weights are the probabilities of choosing these alternatives. From the expression above, it is clear that the signal of the estimated coefficients may differ from the signal of the corresponding marginal effects.

1.5. Results and Discussion

In this section, we will present the results of the estimation of our land-use model and discuss the key findings, comparing them with the theoretical predictions and to other empirical studies. In this way, subsection 5.1 briefly presents the results of the parametric multinomial logit and subsection 5.2 shows the results of the locally weighted estimation.

1.5.1. Standard Multinomial Logit Estimates

Initially, we chose to use five categories of land use in the empirical model: unimproved, residential, commercial, multifamily residential and mixed use. These categories correspond to 97.51% of the total number of observations. We opted to merge the industrial category with commercial ones, since the number of industrial parcels is not very representative (only 0.24% of total) and most of the manufactures in Recife are of light aspect, resembling commercial establishments. The parcels classified as “Special Classes” were not used in the estimation, since they are characterized by generic uses and, in most cases, are not subject to market rules (such as public buildings, hospitals and churches). Thus, Table 2 presents the marginal effects of the multinomial logit model with five categories of urban land use. Additionally, the results of the estimated coefficients (with unimproved use normalized to zero, as a base category) are available in table 3A of the Appendix. The model was estimated by the maximum likelihood method and all the explanatory variables are in logarithmic format. Each column of Table 2 represents a specific category.

The figures in Table 2 indicate that the greater the distance to the Central Business District (CBD), the greater the probability of unimproved parcels, and the lower the likelihood of commercial, mixed use and residential land use. That is, both households and businesses are attracted by the degree of lot centrality. More specifically, a 1% increase in lot distance to CBD reduces the probability of residential use by 5.76% and reduces de probability of commercial use by 8.59%. However, the use of lots for apartment buildings is positively affected by distance to the CBD, a counterintuitive result in face of the urban economics theory, as discussed in subsection 2.2. In relation to transport infrastructures (main avenues, highways and subway stations), we note that mixed use and commercial activity are attracted by these facilities, while residential is discouraged. The multifamily residential use is more common near main avenues and to highways, which may be a consequence of the high land value in these places that incentive capital-intensive buildings.

The numbers of column (5) indicates that mixed use is more frequent in the proximity of the CBD and transport facilities, like the commercial parcels. Although this result is counterintuitive in relation to the theoretical motivation for mixed use (low agglomeration forces and high commuting costs), this suggests that the commercial side of mixed use is more preponderant than the residential side in determining the geographical location of the lot. Regarding natural amenities, a variety of patterns are observed: open areas (which include parks, green areas and squares) attract both types of residential use. However, the proximity to the beach and the river encourages commercial use, multifamily residential use, mixed use and the unimprovement of lots.

Table 2 – The determinants of urban land use: standard Multinomial Logit marginal effects

	(1)	(2)	(3)	(4)	(5)
	Unimproved	Residential	Commercial	R. Multifamily	Mixed Use
Dist. CBD	0.1454***	-0.0576***	-0.0859***	0.0064***	-0.0082***
	(0.0034)	(0.0045)	(0.0021)	(0.0020)	(0.0006)
Area	0.0388***	-0.1408***	0.0440***	0.0533***	0.0047***
	(0.0013)	(0.0019)	(0.0009)	(0.0008)	(0.0003)
Dist. Roadways	-0.0117***	0.0443***	-0.0225***	-0.0081***	-0.0019***
	(0.0014)	(0.0021)	(0.0011)	(0.0010)	(0.0004)
Dist. Main Avenues	0.0195***	0.0019*	-0.0141***	-0.0056***	-0.0017***
	(0.0010)	(0.0012)	(0.0005)	(0.0005)	(0.0002)
Dist. Subway	-0.0059***	0.0116***	-0.0145***	0.0101***	-0.0018***
	(0.0013)	(0.0017)	(0.0007)	(0.0008)	(0.0002)
Dist. Open Areas	0.0424***	-0.0355***	0.0024***	-0.0089***	-0.0005*
	(0.0013)	(0.0017)	(0.0008)	(0.0007)	(0.0002)
Dist. Beach	-0.0602***	0.1377***	-0.0341***	-0.0384***	-0.0051***
	(0.0018)	(0.0027)	(0.0011)	(0.0010)	(0.0004)
Dist. *Capibaribe* River	-0.0529***	0.0682***	-0.0063***	-0.0075***	-0.0015***
	(0.0012)	(0.0017)	(0.0009)	(0.0008)	(0.0003)
N. Observations	95708				
Log-Likelihood	-70884,156				
McFadden R^2	0.1942				

Note: *** p<0.01, ** p <0.05, * p <0.1. The standard errors are in parentheses. The outcome variables are the categories of urban land use: Unimproved, Residential, Commercial, Residential – Multifamily or Mixed Use. The explanatory variables are in logarithm format and the model were estimated by maximum likelihood.

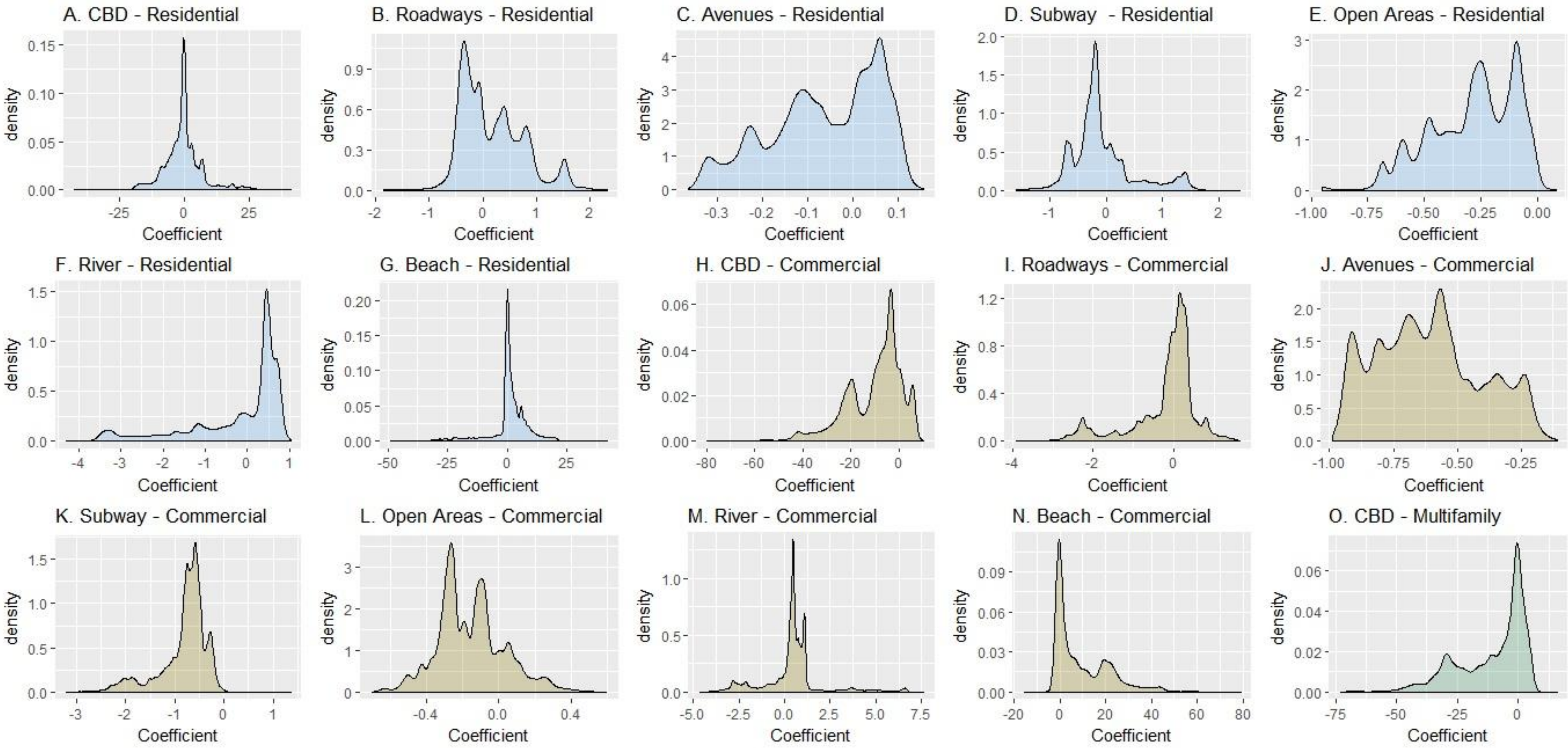
Like the traditional regression framework, a shortcoming of the parametric multinomial logit is that the estimated parameters represent only the mean values. Thus, the estimation process implicitly assumes that all parcels of the city are affected in the same manner by the different urban attributes, regardless of their geographic location. That is, it imposes a single parametric and linear specification to explain all local relationships. This approach is too restrictive in an intra-urban context, where it is possible that there is strong spatial variation (McMillen and McDonald, 2004; Wang et al., 2011). In the next subsection, we will present the locally weighted multinomial logit estimation, a less arbitrary approach that allows parameters to vary smoothly spatially.

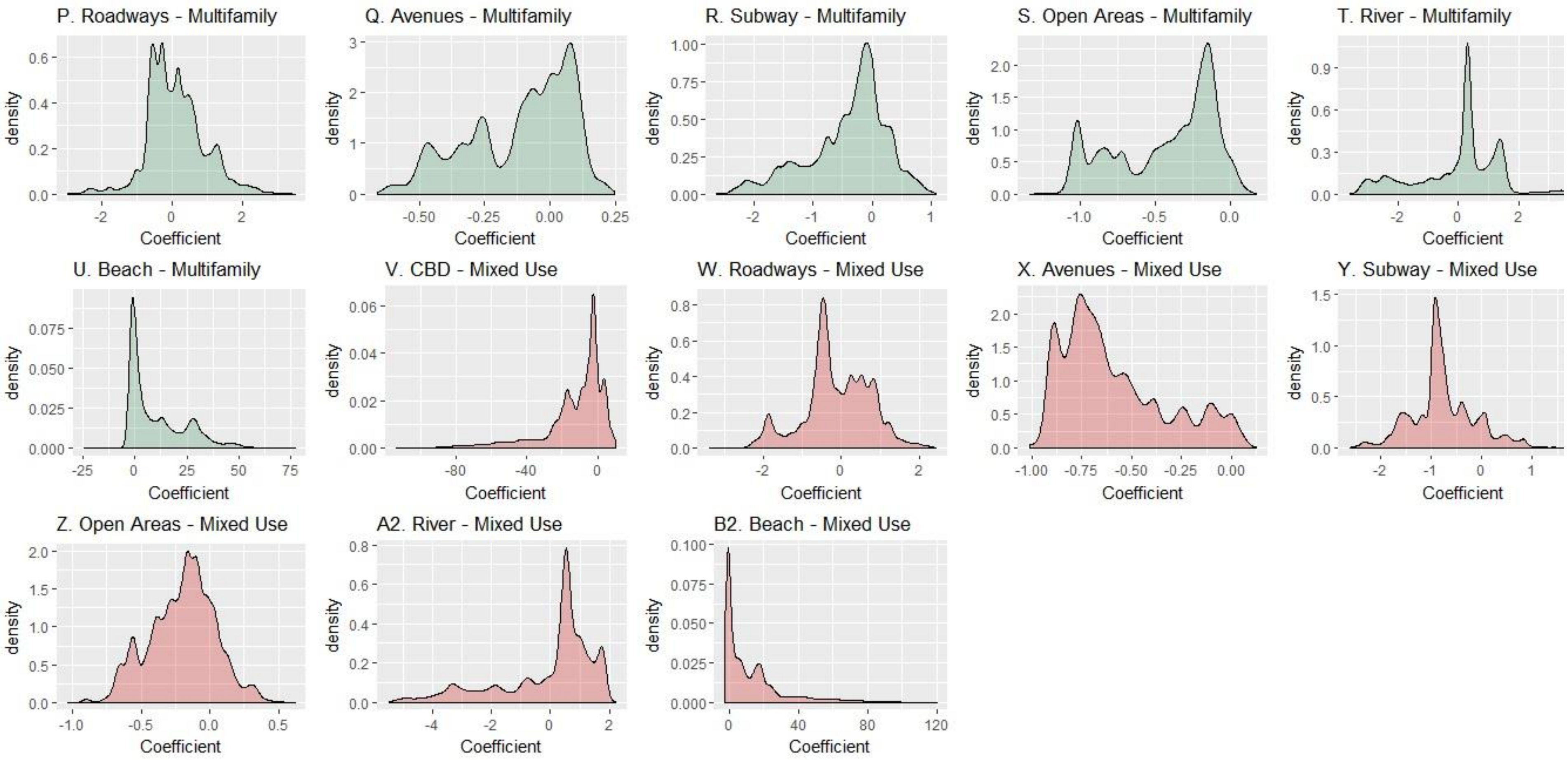
1.5.2. LW Multinomial Logit Estimates

In order to obtain a set of coefficients for each particular observation, we estimate a Locally Weighted Multinomial Logit model through the log likelihood function given by (2). We used a tri-cube kernel weight function and, based on the Cross-Validation method, we chose a 40-percent Window Size. In addition, like McMillen and Soppelsa (2015), we employ the adaptive decision tree approach (Loader, 1999) to estimate the model in a smaller number of target points, which significantly reduces estimation time. Thus, Figure 2 shows the Kernel density functions of the distribution of estimated coefficients, considering the unimproved use as the base category.

Initially, Figure 2 shows a strong heterogeneity in the estimated coefficients, so that in most distributions, there are both positive and negative values. For example, it is noted that the proximity to the CBD may attract or displace any category of land use, indicating that urban areas are much more complex than the relationships established in the monocentric model. In this sense, there are few situations where each specific Kernel density contains only coefficients with a common signal. One such case is the distribution of the coefficients related to the main avenues and to commercial land use (density “J”), where all coefficients take negative values. This reveals that proximity to major avenues encourages commercial use in all parcels of the city, without exception. Furthermore, almost all the coefficients that measure the relationship between public open areas and residential land use (both single and multifamily, density “E” and “S”) have a negative sign, revealing that this type of amenity strongly support the residential development.

Figure 2 – Kernel Densities of Coefficient Distribution, LW Multinomial Logit.





Note: The model is estimated by the locally weighted maximum likelihood using a 422 target points and a tri-cube kernel function with a 40 percent window size. The number of observations is 95,708. The outcome variables are the categories of urban land use: Unimproved, Residential, Commercial, Residential – Multifamily and Mixed Use. Where the Unimproved use is defined as the base category. All the explanatory variables are in logarithm format. The blue distribution represents the coefficients associated to residential land use, the yellow represents the coefficient associated to commercial land use, the green represents the coefficients associated to residential multifamily use, and the red distribution, those related to the mixed use.

Another interesting aspect of the Kernel densities of Figure 2 is that in each of them there is a limited set of parcels that develop similarly among each other but diverges from the complete set, generating the shape with peaks and valleys. For example, we observe that the distribution of coefficients related to proximity to the beach (density "U") indicates that this natural amenity strongly encourages some parcels for multifamily use (the magnitude of the coefficients reaches up to -24.5) but exerts little or no influence in the large part of the lots used for this purpose.

In order to facilitate the interpretation of the results and to identify a general pattern of urban land use, Table 3 presents the mean, standard deviation, maximum and minimum of the marginal effects associated with each coefficient[5].

Table 3 – The determinants of urban land use: LWR Multinomial Logit marginal effects

	(1)	(2)	(3)	(4)	(5)
	Unimproved	Residential	Commercial	R. Multifamily	Mixed Use
Dist. CBD	-0.0707	0.9549	-0.3685	-0.4557	-0.0599
	(1.0066)	(1.3943)	(0.4731)	(0.9758)	(0.1171)
	[-6.70, 16.22]	[-4.42, 9.89]	[-9.99, 2.21]	[-9.32, 1.21]	[-2.41, 0.25]
Area	0.0319	-0.1446	0.0450	0.0631	0.0046
	(0.0425)	(0.0956)	(0.0467)	(0.0692)	(0.0059)
	[-0.21, 0.19]	[-0.47, 0.02]	[-0.12, 0.32]	[-0.01,0.42]	[-0.032, 0.07]
Dist. Roadways	0.0052	0.0395	-0.0460	0.0145	-0.0027
	(0.0494)	(0.1229)	(0.1051)	(0.0622)	(0.0119)
	[-0.58, 0.49]	[-0.50, 0.55]	[-0.53, 0.19]	[-0.41, 0.75]	[-0.15, 0.24]
Dist. Main Avenues	0.0134	0.0147	-0.0262	0.0010	-0.0029
	(0.0181)	(0.0279)	(0.0262)	(0.0116)	(0.0041)
	[-0.01, 0.17]	[-0.06, 0.16]	[-0.22, 0]	[-0.06, 0.21]	[-0.06, 0.01]
Dist. Subway	0.0022	0.0477	-0.0330	-0.0114	-0.0055
	(0.0728)	(0.0982)	(0.0398)	(0.0454)	(0.0089)
	[-0.48, 0.25]	[-0.15, 0.49]	[-0.50, 0.13]	[-0.42, 0.31]	[-0.10, 0.03]
Dist. Open Areas	0.0311	-0.0265	0.0073	-0.0111	-0.0009
	(0.0368)	(0.0359)	(0.0144)	(0.0273)	(0.0055)
	[-0.02, 0.20]	[-0.18, 0.12]	[-0.08, 0.17]	[-0.21, 0.07]	[-0.06, 0.02]
Dist. Beach	0.0580	-0.8201	0.2428	0.4725	0.0524
	(1.1330)	(1.6179)	(0.4771)	(1.0036)	(0.1216)
	[-15.97, 8.38]	[-10.04, 4.28]	[-1.78, 9.91]	[-1.90, 9.71]	[-0.13, 2.55]
Dist. Capibaribe River	0.0196	-0.0525	0.0328	-0.0003	-0.0002
	(0.1361)	(0.1472)	(0.0969)	(0.0600)	(0.0070)
	[-0.55, 0.79]	[-0.92, 0.25]	[-0.27, 1.05]	[-0.66, 0.54]	[-0.17, 0.04]

[5] Before interpreting the results of Table 3, note that, unlike the parametric models, it is not clear how to get an aggregate measure of statistical tests for the average marginal effects. Issues regarding statistical inference still need to be improved by the theoretical LWR literature.

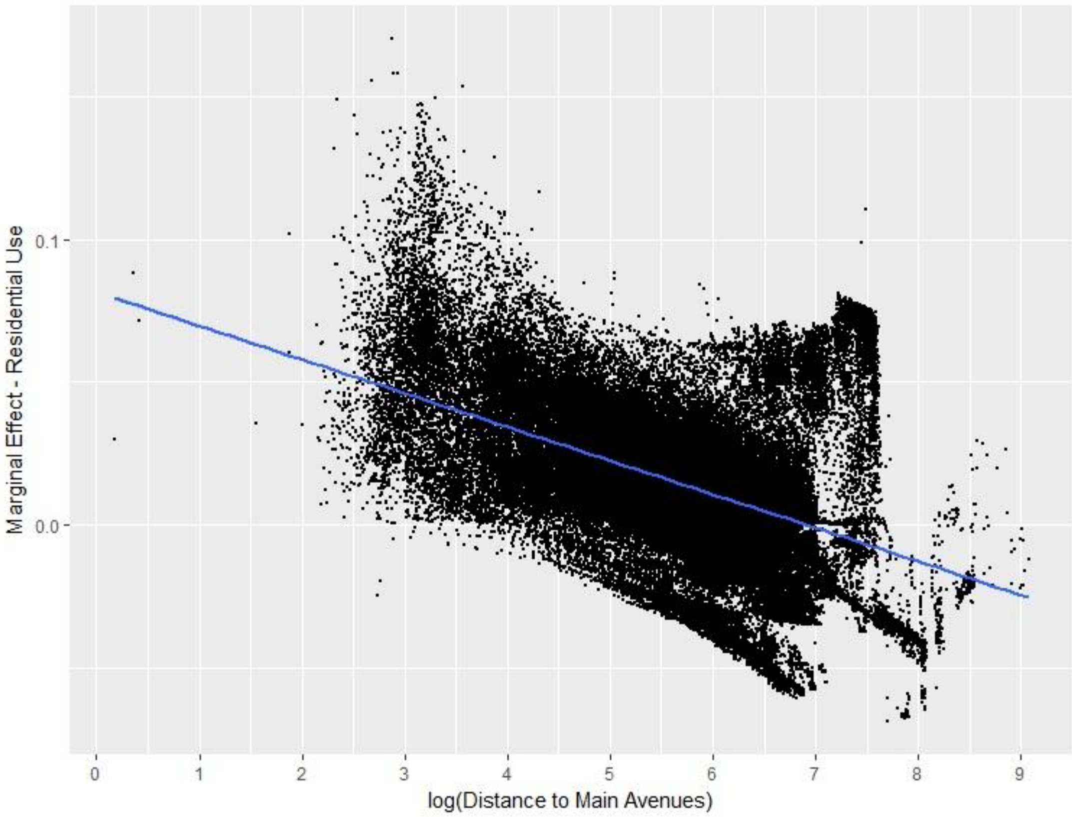
Note: The model is estimated by the locally weighted maximum likelihood using a 422 target points and a tri-cube kernel function with a 40 percent window size. The number of observations is 95,708. The outcome variables are the categories of urban land use: Unimproved, Residential, Commercial, Residential – Multifamily and Mixed Use. All the explanatory variables are in logarithm format. The standard deviations are in parentheses and the minimum and maximum are in brackets.

From Table 3, it can be observed that the greater the distance to the CBD, the smaller - on average - the probability of the unimproved land. This indicates that there is, in some way, a concentration of empty or vacant spaces near the historic center. Nadalin and Igliori (2016) found a very similar result for the case of São Paulo. According to the authors, a feasible explanation for the proliferation of empty uses in the historic center is that the place has lost some of its attractiveness as the city became more polycentric and thus began to receive a smaller share of public investments. In addition, in these places, there is greater incentive for landowners to wait for higher prices rather than sell or rent their properties, since any investment in renewing the historic center tends to prompt rapid local gentrification (Brueckner and Rosenthal, 2009). In relation to columns (2) and (3), it can be identified that, evem though CBD can both attract and displace commercial and residential land use, commercial development tends to be higher in areas close to the center (negative mean), while the residential use tends to be lower in these areas (positive mean). The strong variability of the marginal effects associated with residential use (ranging from -4.42 to 9.87) indicates that, in fact, households have heterogeneous preferences regarding the accessibility-space choice (Duranton and Puga, 2015). Column (4) of Table 3 shows that, on average, the probability of land use for apartment buildings is negatively correlated with distance to CBD, a result predicted by urban economics. Locations closer to the city center tend to have more expensive lands, which encourages capital-intensive (rather than land intensive) constructions (Fujita, 1989; Ahlfeldt and McMillen, 2017). Mixed land use also tends to occur near the CBD (column 5), although there are parcels that are positively influenced. This result may be explained by the fact that a great part of the buildings of the historic center of Recife were built in the eighteenth and nineteenth centuries, a period characterized by high commuting costs and by a public administration that employed urban planning traditions that prioritized mixed land use. The historical characteristic of the city ends up offsetting the agglomeration forces that exist in the center that could encourage the conversion of mixed uses to purely commercial uses. This evidence is similar to that of McMillen and McDonald (1999), which analyzed the city of Chicago in 1920 and showed that mixed use was more common near the CBD and near the commuter train stations.

In relation to transport infrastructures, Table 3 presents a series of interesting results. First, the proximity to highways and main avenues affects very similarly the different categories of land use. On average, these facilities attract the business sector and deviate households (both standard and multifamily residences). Similarly to Figure 2, there are no cases where the proximity to the main avenues discourages commercial use, since the marginal effects only take negative or zero values (ranging from -0.22 to 0). As discussed earlier, highways and avenues are key variables for the locational choice of firms, since these infrastructures facilitate the flow of products, the commuting of employees and allow easier access to the consumer market (O`Sullivan, 2012, Billings and Johnson, 2016). That is, these infrastructures attract large population flows and generate agglomerative gains. The empirical literature of urban land use also points out that the greater proximity to highways, freeways, highways and main avenues encourage both the commercial and the industrial development (Braimoh and Onishi, 2007; Wang et al., 2011; Jacob and McMillen, 2015). From column (1), we observe that parcels distant from these transport infrastructures tend to have a greater probability to be unimproved. From column (5), we note that mixed use is more common near these facilities, reinforcing the idea that the commercial side of mixed use is more decisive than the residential side in determining the geographic location of the lot. In relation to residential sector, the results of column (2) and column (4) indicate that, on average, households choose to live in areas further away from highways and main avenues, an indication that negative externalities (noise, pollution and crowding) overlap the benefits related to greater accessibility.

To understand better this effect, Figure 3 plots the correlation between the distance to the main avenues and the corresponding marginal effects of residential use (each point corresponds to a particular lot). Note that there is a strong non-linearity: as the distance from the avenues increases, the marginal effects decrease and reach negative values, indicating that the households begin to appreciate these transportation facilities. This behavior is consistent with the idea that the negative externalities generated by the main avenues are only concentrated in parcels very close to this infrastructure and support the empirical studies showing that local residents are commonly opposed to the installation of transport infrastructures in their immediate proximity (Ahlfeldt and Maennig, 2015).

Figure 3 - Correlation Between Marginal Effect of Residential Use and Distance to Main Avenues.



Note: The Marginal Effects were extracted from the specification described in Table 3, the distance to main avenues are in logarithm format and each point corresponds to a particular lot. The blue line corresponds to the regression line.

The last three lines of table 3 shows the relations between natural amenities (public open areas, *Capibaribe* River and the Beach) and the five land use categories. Looking at the average marginal effect, a general result is easily identified: all three amenities tend to attract the residential sector and discourage commercial activities and the unimprovement of lots. The only counterintuitive result is that the beach reduces – on average - the probability of multifamily residential use. A simple visual inspection of the city's seashore reveals that it is replete of apartment buildings (Figure 1). The large variability of the marginal effects (ranging from -1.9 to 9.71) and the fact that the seashore concentrates only a small proportion of the total number of apartments in the city can explain this result. This evidence ratifies the urban economic theory that points to natural amenities as key factors for the households' locational choices, influencing the spatial allocation of land use in cities (Brueckner et al., 1999). In addition, the empirical literature of urban land use finds results similar to ours. For example, Braimoh and Onishi (2007) show that residential development tends to be closer to water bodies, and McMillen and Soppelsa (2015) find that proximity to parks and to Lake Michigan increases the probability of residential land use. The hedonic prices applications also evidenced the importance of natural amenities: Seabra et al. (2016) shows that the proximity to the *Capibaribe* River and to the Beach appreciate the price of housing in Recife.

To conclude, the figures in Table 3 show a general pattern of urban land use: while transport infrastructures are critical to business activity, natural amenities play a key role in attracting the residential sector. Note that these results diverge from the estimations of Table 2, showing that is extremely important to consider the wide local coefficient variation and the respective non-linearities. In the next section, we will analyze the predictive capacity of the empirical models.

1.6. Model Predictive Performance

In order to evaluate and compare the predictive performances of the parametric and the locally weighted Multinomial Logit, Table 4 shows the land use predictions of both approaches, comparing them with the actual land use. As a starting point, the first number of each cell in the table shows the predictions constructed under the assumption of random land use, using the sample proportions. For example, note that 11,074 of the lots are unimproved, representing 11.57% of the total, so if we randomly predict the land uses, we would predict that 0.1157 x 69250 = 8012.6 of the truly residential parcels are unimproved. The second number of each cell represents the prediction of the standard multinomial logit and the third number shows the prediction obtained with the LW multinomial logit. Thus, from the first cell, it can be observed that the parametric model accurately predicts 851 unimproved parcels, while the nonparametric model correctly predict 1947 of these. The diagonal elements are of greater interest, since they represent the number of correct predictions of each particular empirical model.

Table 4 – Predictive Performance.

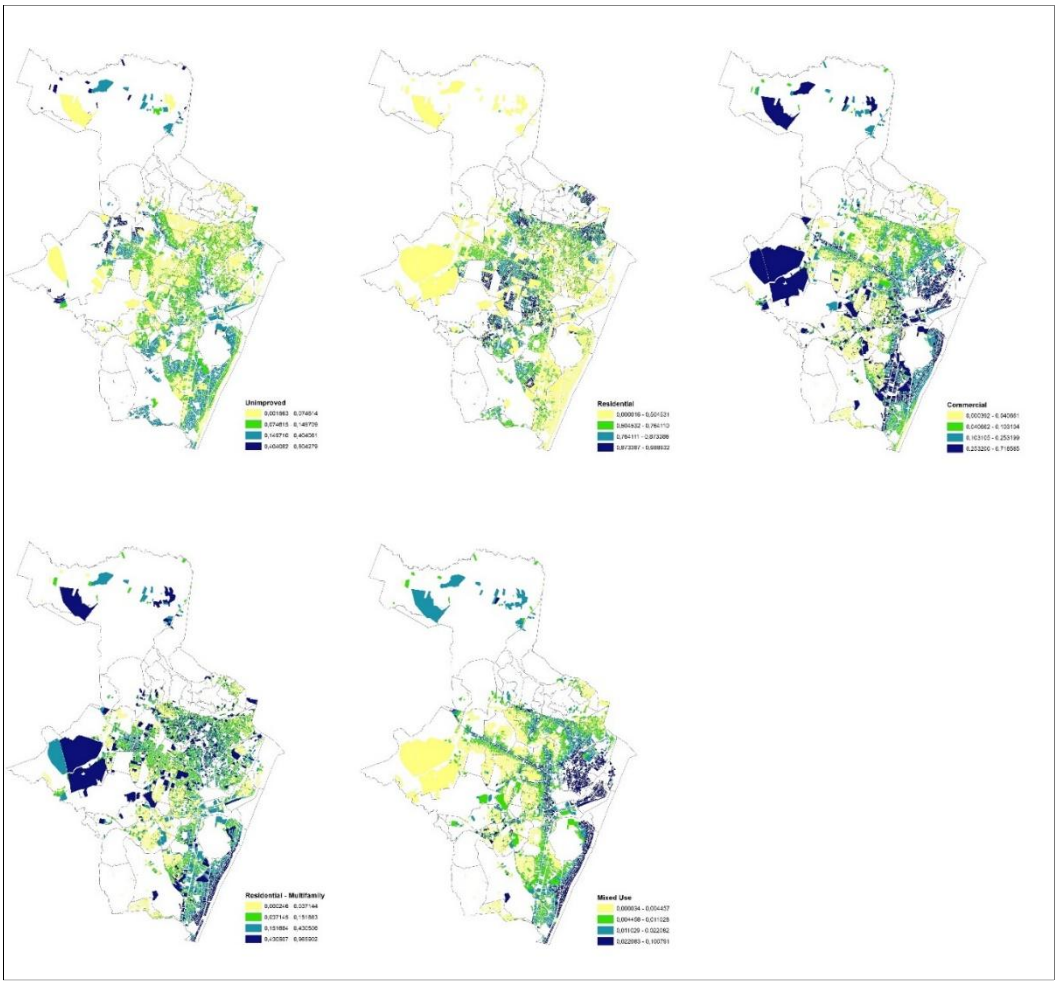
Actual Use	Predicted Land Use					
	Unimproved	Residential	Commercial	R. Mult.	Mixed	Total
Unimproved	1281.3	8012.6	907.6	760.2	112.2	11074
	851	9318	306	599	0	
	1947	8232	330	565	0	
Residential	8012.6	50106.2	5675.6	4753.8	701.8	69250
	610	67413	695	532	0	
	859	67080	733	578	0	
Commercial	907.6	5675.6	642.9	538.5	79.5	7844
	88	5250	1822	684	0	
	138	4694	2410	602	0	
R. Multifamily	760.2	4753.8	538.5	451.0	66.6	6570
	56	4696	232	1586	0	
	91	4401	254	1824	0	
Mixed Use	112.23	701.84	79.49	66.58	9.83	970
	2	642	228	98	0	
	3	616	252	99	0	
Total	11074	69250	7844	6570	970	95708
	1607	87319	3283	3499	0	
	3038	85023	3979	3668	0	

Note: The first number of each cell presents the random prediction (based in sample proportions), followed by the parametric prediction and LWR predictions.

Thus, the diagonal numbers of Table 4 reveals that the LW Multinomial Logit has a better predictive capacity compared with the other two approaches, since it can correctly predict a greater amount of parcels, with the exception of residential land use. The best predictive performance of the locally weighted models is also evidenced in other studies (For example, in McMillen and McDonald, 2004). The numbers of Table 4 reveal that it is extremely difficult to predict the allocation of land for mixed use with the empirical models, which may be a consequence of the low proportion of lots in this category (only 1% of the total).

To conclude, Figure 4 plots the estimated LW probabilities of each land use category. Through a simple visual inspection, it is possible to note the existence of a pattern of spatial clusters, suggesting that geographically close parcels tend to develop very similarly. Additionally, we can see results that resemble the current land use (Figure 1). Like the greater probability of multifamily residential use on the seashore, the concentration of commercial activities in the northeast (near to the city CBD) and the predominance of unimproved lots in suburban areas.

Figure 4 – Predicted Probabilities from LW Multinomial Logit Model.



Note: The ranges were based on natural breaks (Jenks) intervals

1.7. Conclusions

In the last decades, the Latin American cities experienced intense modifications in their spatial configurations, including the increase of verticalization, urban sprawl, slum formation and the formation of polycentric structures. Faced with these new tendencies and in response to urban problems, public policies towards cities started to gain greater priority from central governments. In the Brazilian case, several policies were adopted, including the creation of a federal law that establishes general guidelines for the subdivision of lots and zoning (Federal Law 6,766/79), the creation of the Ministry of Cities in 2003 and the implementation of a housing subsidies policy of approximately R$ 34 billion.

In this context of city growth and greater awareness of urban challenges, it becomes important to understand how the interactions work between households, business and the urban space. Thus, the objective of this paper was to explain the patterns of urban land use for the city of Recife, a large Brazilian metropolis. More specifically, with a novelty microdata set of 98,198 individual lots, we investigated the role of transport infrastructures, lot centrality and the availability of natural amenities on the explanation of different categories of urban land use. Note that like other cities in developing countries, Recife has a large informal housing market. This makes our exercise particularly interesting since it allows us to understand to what extent the competitive market is able to explain the allocation of land use in a heavily informal city.

Through a locally weighted Multinomial Logit, we detected that there are considerable local spatial effects, indicating that parametric approaches are inadequate to explain land use in an intra-urban context. Additionally, we also noted that the LWR approach has a better predictive performance compared to parametric counterpart. Although there are significant local spatial variations, we identified some general patterns of land use: transport infrastructures (which include main avenues, highways and subway stations) tend to attract commercial use, while natural amenities (including the public open spaces, the beach and the river) tend to encourage residential development. In addition, our results indicate that the probability of unimproved use is greater in areas further away from natural amenities and transport facilities and in areas close to the CBD. It has also been shown that, on average, verticalization is greater in areas close to the CBD and to natural amenities,although it is completely expected that the intensity of use be higher in more valuated areas, the proliferation of uses that generate negative externalities (such as multifamily residences) near rivers and the seashore indicates that the zoning ordinances of the city was ineffective in encouraging the private internalization of these externalities.

Finally, another interesting result is that mixed use occurs in areas near transportation facilities and the CBD, indicating that the commercial side of mixed use is more decisive than the residential side in determining the geographic location of the lot.

In general, our results indicate that the patterns of urban land use in Recife do not differ significantly from those of North American cities (McMillen and Soppelsa, 2015; Wang et al., 2011), and that a great part of the theoretical conclusions of urban economics are empirically corroborated. In practical terms, our results can be useful to improve the city urban planning and help formulate public policies for urban land use. First, the choice of locations for the implementation of social housing programs should prioritize areas that have a natural vocation for residential development. The definition of urban zoning must also respect the equilibrium obtained through the competitive market - which generates a clear segregation of uses - in order to avoid conflicts and the increase of unimproved parcels. In addition, the decision to build transport infrastructures should also consider the potential changes in the land use of the vicinity areas. Finally, the high number of vacant lots in the city (especially in the historic center) indicates that there is still a gap for urban growth, without the need to increase verticalization or sprawl.

1.8. References

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1.9. Appendix A – First Chapter

Table 1A – List of Recife Main Avenues

Av. Eng. Abdias de Carvalho	Av. Herculano Bandeira	R. Sebastião Malta Arcoverde
Av. Gov. Agamenon Magalhães	R. João Ivo da Silva	R. Real da Torre
Av. Visc. de Albuquerque	R. Joaquim Nabuco	Av. Recife
R. Alegre	Av. Joaquim Ribeiro	R. Regeneração
R. Augusto Calheiros	R. José Bonifácio	Est dos Remédios
R. Cônego Barata	R. José Fernandes de Souza	R. Ribeiro de Brito
Cor. do Bartolomeu	R. José Osorio	R. Ricardo Salazar
Av. Beberibe	Av. Dr. José Rufino	R. Padre Roma
R. Benfica	R. Arquiteto Luiz Nunes	R. Tito Rosas
R. São Bento	Av. Mascarenhas de Moraes	R. Prof. Trajano de Mendonça
R. Bomba do Hemetério	R. Nicolau Pereira	Est Velha de Agua Fria
Av. Caxangá	R. Dona Olegarinha	Av. Prof. Estevão Francisco da Costa
R. Cosme Viana	R. Olivia Menelau	R. Quitério Inácio de Mello
Av. Eng. Domingos Ferreira	R. Visc. de Pelotas	Av. Boa Viagem
R. Ernesto de Paula Santos	R. Dr. Samuel Lins	Av. Antônio de Góes
R. Córrego do Euclides	Av. Gal San Martin	R. Visc. de Jequitinhonha
R. Falcão de Lacerda	R. Santos Araújo	Av. Norte Miguel Arraes de Alencar
R. da Harmonia	R. São Sebastião	

Note: classified as arterial routes by city's transit agency.

Table 2A – List of the Thirty Largest Public Open Spaces

Praça Dr. Adolfo Cirne (Square)	Derby (Square)
Cemitério de Santo Amaro (Square)	Tertuliano Feitosa (Square)
Sítio da Trindade (Park)	Campo do Bueirão (Square)
Santana (Park)	Beira-Rio (Park)
Treze de Maio (Park)	Chesf (Square)
Caiçara (Park)	Camilo Pereira Carneiro (Square)
Engenheiro Abdias de Carvalho (Park)	Memorial Arco-Verde (Park)
Lagoa do Araçá (Green Area)	Dona Lindu (Park)
Ministro Salgado Filho (Park)	Exposição do Cordeiro (Park)
Forte do Arraial Novo do Bom Jesus (Park)	Jaqueira (Park)
Rua da Aurora (Green Area)	Maria do Carmo Araújo (Green Area)
Dr. Arnaldo Assunção (Park)	Jardim Botânico (Green Area)
João Francisco Lisboa (Square)	Dois Irmãos (Green Area)
Novaes Filho (Square)	Guabiraba (Green Area)
Robert Kennedy (Park)	Pau-Ferro (Green Area)

.

Table 3A – Standard Multinomial Logit Coefficients.

	(1)	(2)	(3)	(4)
	Residential	Commercial	R. Multifamily	Mixed Use
Intercept	9.5402***	31.4409***	9.5248***	29.0399***
	(0.413)	(0.634)	(0.666)	(1.352)
Dist. CBD	-1.4926***	-3.1747***	-1.2717***	-2.8231***
	(0.038)	(0.056)	(0.057)	(0.096)
Area	-0.5550***	0.5196***	0.8650***	0.4222***
	(0.015)	(0.020)	(0.019)	(0.039)
Dist. Roadways	0.1695***	-0.3464***	-0.0758***	-0.2110***
	(0.015)	(0.027)	(0.026)	(0.070)
Dist. Main Avenues	-0.1887***	-0.4787***	-0.3226***	-0.4875***
	(0.011)	(0.015)	(0.016)	(0.032)
Dist. Subway	0.0724***	-0.2478***	0.3145***	-0.2408***
	(0.014)	(0.020)	(0.023)	(0.043)
Dist. Open Areas	-0.4591***	-0.3645***	-0.6216***	-0.4978***
	(0.015)	(0.022)	(0.020)	(0.044)
Dist. Beach	0.7605***	-0.1074***	-0.3077***	-0.2795***
	(0.020)	(0.026)	(0.024)	(0.052)
Dist. *Capibaribe* River	0.6022***	0.3881***	0.3409***	0.2594***
	(0.013)	(0.023)	(0.022)	(0.047)
N. Observations	95,708			
Log-Likelihood	-70884,156			
McFadden R^2	0.1942			

Note: *** p<0.01, ** p <0.05, * p <0.1. The standard errors are in parentheses. The outcome variables are the categories of urban land use: Unimproved, Residential, Commercial or Residential - Multifamily. Where the Unimproved use is defined as the base category. The explanatory variables are in logarithm format and the model were estimated by maximum likelihood.

Table 4A – LW Multinomial Logit Coefficients: Means and Standard Deviations.

	(1)	(2)	(3)	(4)
	Residential	Commercial	R. Multifamily	Mixed Use
Intercept	3.6385	11.7964	-0.3268	6.2948
	(11.577)	(18.904)	(17.988)	(21.297)
Dist. CBD	-0.5087	-9.9928	-10.8193	-12.6211
	(7.605)	(11.508)	(14.058)	(17.890)
Area	-0.7096	0.5128	0.8300	0.3360
	(0.387)	(0.466)	(0.314)	(0.367)
Dist. Roadways	0.1675	-0.2240	0.0988	-0.1220
	(0.576)	(0.826)	(0.799)	(0.861)
Dist. Main Avenues	-0.0619	-0.6108	-0.1141	-0.5874
	(0.120)	(0.204)	(0.198)	(0.263)
Dist. Subway	-0.0778	-0.8184	-0.4026	-0.7606
	(0.540)	(0.489)	(0.656)	(0.639)
Dist. Open Areas	-0.2846	-0.1549	-0.4313	-0.1988
	(0.185)	(0.186)	(0.331)	(0.237)
Dist. Beach	1.1919	9.2561	10.9258	12.7446
	(8.176)	(11.885)	(14.303)	(18.478)
Dist. *Capibaribe* River	-0.1963	0.4322	0.0026	-0.0857
	(1.123)	(1.605)	(1.365)	(1.623)
N. Observations	95708			
Log-Likelihood	-66220.48			

Note: The model is estimated by the locally weighted maximum likelihood using a 422 target points and a tri-cube kernel function with a 40 percent window size. The outcome variables are the categories of urban land use: Unimproved, Residential, Commercial, Residential – Multifamily and Mixed Use. Where the Unimproved use is defined as the base category. All the explanatory variables are in logarithm format. The table shows the means of the coefficients and the standard deviation (in parenthesis).

Chapter 2 - The Impact of Zoning Ordinances on the Housing Market in Developing Counties: Evidence from Brazilian Municipalities

2.1. Introduction

In recent decades, a sharp increase in population concentration took place in major Brazilian metropolitan areas, continuing the rapid urbanization process that has been occurring in Brazil over the last century. For example, the population density of Brazilian urban areas grew 16.65% between 2000 and 2010. In new metropolitan areas such as Brasilia and Goiania, the population density increased by 23.7% and 22.7%, respectively (Brazilian Demographic Census 2010). Due to higher population density around cities and macroeconomic stability, which increased housing credit, the Brazilian real estate market underwent a price boom. According to the Fipe Zap index, which measures the selling price of properties and rentals in main Brazilian cities on a monthly basis, from 2010 to 2015, housing prices increased 223.5% in Sao Paulo, whereas rents increased at a rate of 100.8%. This variation was significantly higher than inflation in the same period (which reached 54.6%, measured by IPCA, the official Brazilian consumer price index).

In addition to demand-related factors, supply-related factors such as the higher cost of real estate development and the restrictions in urban land use may have also a connection with this great appreciation in the Brazilian real estate market. One of the main ways to regulate and restrict urban land use in Brazil is through zoning ordinances. Zoning ordinances are usually established by local governments and can regulate both the type of use (establishing residential areas, commercial areas, industrial or mixed) and the intensity of use, to restrict the size, the weight of buildings or the flood-area ratio (McDonald and McMillen 2012). Note that, in addition to traditional zoning, there are other regulations that can be used to restrict urban land use, such as rules governing septic systems, wetland regulations, green area requirements, and others (Fischel, 2004; Glaeser and Ward, 2009).

In Brazil, local governments (municipalities) are in charge of the adoption and the management of land use controls. These instruments have gained greater relevance in the late 70s and thus became regulated by federal laws. In 1979, a federal law that regulates urban land parceling (Federal Law 6766/79) came into force. A key feature of this law is that it establishes a minimum lot of 125 square meters and, in addition, delegates to municipalities the autonomy to set their own guidelines for the allotment and the breakup of land. In 2001, the federal government created the "*Estatuto da Cidade*" (Federal Law 10257/01), a law that established general urban guidelines and the obligation of master plans for a set of municipalities. In this context of federal regulation on urban planning instruments, a large increase in the number of municipalities that have their own zoning ordinance occurred: while in 1978, only 64 municipalities had regulations for the urban land, in 2013 this number rose to 1,724 (Munic 2013). Despite the recent increase in the adoption of these regulations and the vast empirical literature considering the effects of urban regulation, little is known or investigated about their impact on the Brazilian housing market.

According to McDonald and McMillen (2012), the objective of the supporters of the law is to protect the residents' well-being from the adverse effects of industrialization, rapid and intense urbanization and high-impact buildings. Therefore, zoning would be useful to reduce the negative externalities that are associated with certain types of land use or degrees of use. However, by restricting the land use and limiting new developments, urban zoning ends up bringing changes to the land market and consequently may increase the prices of existing buildings and residences (Quigley 2007; McDonald and McMillen 2012). Thus, it creates a great incentive for homeowners to support, defend and claim for zoning laws (or other restriction instruments) with the aim of increasing the price of their houses, even in the absence of actual negative externalities (Quigley 2007; Fischel 2004).

A large number of empirical studies investigates the impact of zoning and land use restrictions on the housing and urban land market. Quigley and Rosenthal (2005), Quigley (2007) and McDonald and McMillen (2012) make a detailed review of these studies. Even with different empirical strategies and analyzing different metropolitan areas in the United States, these papers show a similar conclusion: the degree of land use restrictiveness is positively correlated with property price and negatively correlated with new real estate developments (Quigley and Raphael 2005; Ihlanfeldt 2007, Glaeser and Ward 2009; Zabel and Dalton 2011). Additionally, there is evidence that lot price is higher in more regulated areas (Kok, Monkkonen and Quigley 2014) and that cities with more strict regulation suffer from a greater amplitude in the house price cycle, with more intense booms and bursts (Huang and Tang 2012).

The major empirical challenge of those papers is to solve the potential endogeneity in the relationship between zoning and real estate prices. Typically, to overcome this problem, some studies add a wide range of control variables (Glaeser and Ward 2009 and Kok, Monkkonen and Quigley 2014), add fixed-effects (Zabel and Dalton 2011) or employ an instrumental variable approach (Ihlanfeldt 2007). However, not all omitted variables can be measured or are fixed in time, which weakens the first two approaches. Moreover, it is extremely difficult to find an instrument that satisfies the exclusion restriction. For example, Ihlanfeldt (2007) received criticism for using past demographic characteristics as an instrument to the current level of land use restrictiveness. It turns out that past demographics can be correlated with other current characteristics that affect property price, weakening the exclusion restriction (Glaeser and Ward 2009). In addition to this unresolved empirical difficulty, another limitation of those studies is that they focus mostly on the US housing market. Due to the recent urbanization process, the real estate market of developing economies can behave in a very divergent way (Alterman 2013), so that the conclusions of such papers cannot be easily extended to these countries.

In this study, we will take advantage of the variability of the law between the Brazilian municipalities to assess their impact on the average rental prices and on the real estate market growth. This paper contributes by analyzing the Brazilian case, where, unlike the United States, the urbanization process only took off in the second half of the twentieth century. In addition, this study deals with endogeneity in a different way: rather than trying to solve the problem thoroughly, we will evaluate how robust is our evidence in the presence of endogeneity caused by omitted variables. In this sense, considering our goal of investigating the effect of zoning in average rent prices and in the real estate market growth, we employ a Propensity Score Matching (PSM) estimator and a set of sensibility analysis proposed by Rosenbaum (2002), Inchino et al. (2008) and by Oster (2016). Furthermore, another contribution of our study is to use a measure of rental prices, which is adjusted by the property's characteristics.

Our results reveal that municipalities that adopt zoning as an urban planning policy have an average rental price 5.4% to 6.26% higher than that of similar municipalities that do not adhere to the order. The sensitivity analysis based on Rosenbaum bounds indicates that for our evidence to be invalidated there must be an unobservable variable that increases the odds in zoning adoption by 90% and is strongly correlated with the rental price. Additionally, the sensitivity proposed by Inchino et al. (2008) shows that even if there are important and specific confounders, our result remains valid. The Oster (2016) procedure also points in the same direction: for our ATT to be invalid, the unobservables confounders would need to be twice as important as the observables. Thus, the sensitivity tests point that our evidence is hardly driven by endogeneity. Additionally, we also found that the zoning law is not able to affect the real estate market growth, suggesting that higher rental prices are not simply a reflection of a lower housing supply.

The paper is organized as follows: in Section 2, we discuss the evidence of previous literature that assess the impact of land use restrictions in the perspective of developing economies. In Section 3, we present the institutional background and the origins of zoning in Brazilian municipalities. Section 4 details the empirical strategy. Section 5 describes the data, the covariates and the procedure for constructing a constant quality rental price; Section 6 presents the main results and robustness tests and, finally, Section 7 presents the final remarks.

2.2. Zoning in developing countries: empirical evidence

In developing and more recently urbanized countries, policies that restrict land use or that control the intensity of use have consequences on the housing market that can differ substantially in relation to developed countries' markets (Alterman 2013). Firstly, strict regulations with low enforcement power (which is common in developing countries) may generate a low level of compliance and therefore little influence in land and property prices. In addition, stricter regulations - by increasing the price of real estate development and reducing the supply elasticity of available lots - can cause an additional incentive to low-income households decide to enter in the informal housing market and, thus, there is an potential increase in slums formation and in the number of low quality settlements (Cavalcanti and Da Mata 2013). Furthermore, the motivations for the regulation can also be different. For example, in countries with lower income levels, the land-use restrictions can be used as a way to prevent strong densities, which could generate an overload in the poor urban infrastructure (Brueckner and Lall 2015).

In a study conducted for Argentina, Monkkonen and Ronconi (2013) detected that in municipalities with more stringent regulations, both the level of compliance and the price lots were smaller. Although the negative correlation between the restrictions in urban land use and prices is counterintuitive in the face of the US cities' evidence, the authors argue that, by stimulating the growth of low quality houses and slums, land-use restrictions cause negative externalities in the neighborhood and, therefore, the lots suffer devaluation. In a descriptive study for the Offinso South Municipality, in Ghana, Boamah (2013) shows that the land-use regulations have little practical effect, since most of the developers do not comply with the order. According to the author, this is a common situation in African cities and it happens because the planning laws have a low enforcement level, are poorly adapted to local conditions and are guided by public offices with high levels of corruption. Similar results were found by Arimah and Adeagbo (2000) for a Nigerian city.

Although the low applicability of planning laws is an important point in the debate about the effects of zoning in cities in developing countries, the consequences regarding housing affordability and slum formation have an even greater importance on the well-being of citizens. According to UN-Habitat (2003), about 30% of the world's urban population lives in slums, where, due to the poor condition of public infrastructure and low-quality housing, the quality of life of dwellers is very low. Some studies sought to verify the impact of the land-use restrictions on the growth of the informal housing market, considering the Brazilian case. In this sense, Biderman (2007) argues that the formal housing market is very connected to the informal ones. The author also shows evidence that there has been an increase in informality in Brazilian municipalities that have adopted zoning rules in the 1990s. More recently, Cavalcanti and Da Mata (2013), through a structural general equilibrium model, show that much of the 1980-2010 slum growth in Brazil is connected with rural-urban migration, with an increase in income levels and in the restrictiveness of land-use regulations.

Perceiving that land-use restrictions can increase house prices and, thus, worsen the welfare of poor households, some governments use inclusionary zoning as a way to generate affordable housing. However, despite being well intentioned, this policy does not always attain the desired effects. For example, investigating Brazilian cities, Lall, Wang, and Da Mata (2007) show that in municipalities that have adopted urban policies targeted at reducing the minimum lot size, there was an increase in the formal housing supply and an increase in population growth (via migration). Once the population growth was higher than the formal housing growth, the policy also caused an undesirable effect: an increase in slum formation. This evidence is in line with the work of Schuetz, Meltzer and Been (2011) showing that policies that aim to produce affordable houses through inclusionary zoning can also generate the opposite effect: higher prices and lower production rates.

It is also necessary to recognize that as one of the main instruments for urban planning, land-use restrictions and zoning affect many other aspects of the cities' dynamics, going far beyond the effects on the housing market. In a study for Indian cities, Brueckner and Sridhar (2012) found that municipalities with stricter building height limits have larger spatial sizes. The authors show that a unity increasing in the floor-are ratio (FAR) limit can generate a reduction of approximately 20% of the area of an average city, which generates a yearly savings in commuting costs of about 0.7% of the annual income of a typical household.

As can be seen by the above set of evidence, the majority of empirical studies focuses on assessing the impact of the urban land regulations on indirect outcomes, such as the informal housing market. However, for regulations to affect this type of market, it is first necessary to prove that regulations also affect the prices of the formal housing market. Without this, the previous evidence is not very credible. Regarding the Brazilian case, little is known about the consequences of zoning in relation to rental prices or property prices. As far as we know, the study of Dantas et al. (2018) for the city of Recife (the center of the fourth largest metropolitan urban area of Brazil) is the only exception. More specifically, the authors investigated the price effect of a law that restricts the height of the buildings in a particular area of the city. Through a Geographic Discontinuity Differences-in-Differences strategy, the authors showed that the new ordinance had caused an increase of about 7% in the price of properties located in the area in which the ordinance was implemented. However, since this paper analyzes a municipal policy made for a relatively small area, it lacks external validity. In order to generalize the effect of land-use restrictions in Brazil, it is necessary to investigate the problem in a broader perspective. This study aims to fill this gap.

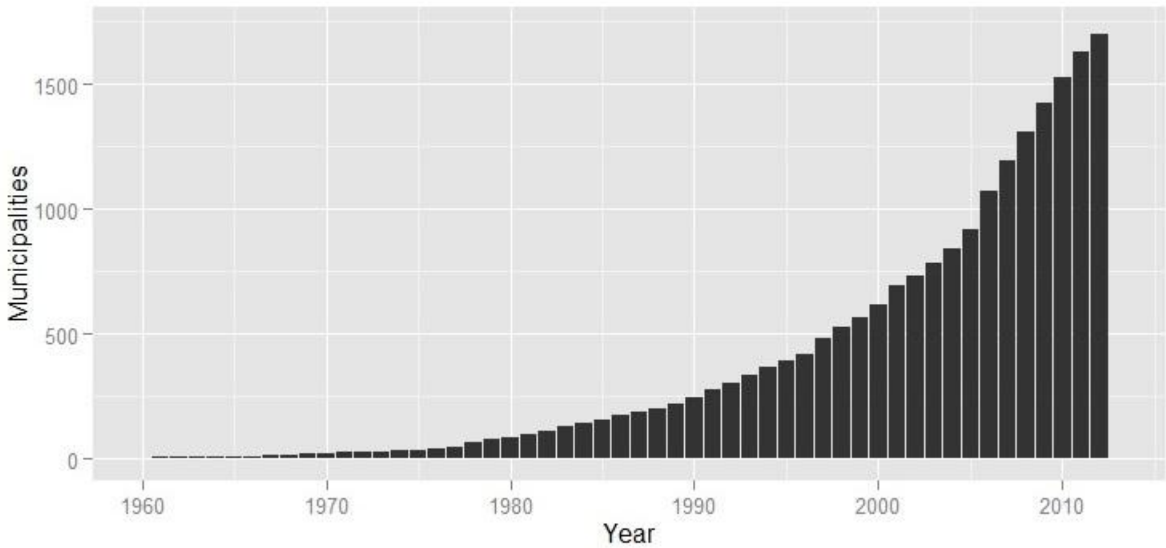
2.3. Zoning and Urban Planning in Brazil: Historical and Institutional Background.

Brazil is a country with relatively recent urbanization: in 1940, only 31.2% of the population lived in urban areas. That number grew throughout the century, rising to 84.3% in 2010 (Census 2010). This rapid urbanization process, linked to the industrialization of the country, brought a number of problems for the Brazilian cities. Not surprisingly, it becomes necessary to introduce some urban planning instruments. Because of its federative character, Brazilian local governments (municipal level) are responsible for creating and establish their own planning laws. It was only at the end of the 1970s that the federal government has assumed a coordinating role in urban issues involving municipalities.

Initially, the first urban regulations used in the country consisted in building codes and limited ordinances. They were deployed in the cities of Rio de Janeiro and Sao Paulo in the late nineteenth century, and, in many cases, resembled zoning rules with a more restricted coverage area. These instruments were demanded by high-income population, who used it as a way to defend their properties’ value against the aggravation of the problems generated by industrialization (Borges 2007). For example, the posture codes of Sao Paulo, created in 1886, prohibited the construction of low-income and working houses in the city's commerce perimeter (Nery Júnior 2013). It was only in the 1930s that urban regulations started being disseminated to other Brazilian cities, such as Recife and Porto Alegre, but were still not comprehensive enough, and thus did not constitute a general rule of zoning.

In the mid-1970s, a spread of general zoning ordinances took place. A milestone is the introduction of the law 7805/72 in São Paulo city, which, among other things, covered the whole territory of the municipality and instituted exclusively residential areas and commerce corridors (Nery Júnior 2013). Figure 1 shows the evolution in the number of municipalities that adopted their own zoning ordinance: in 1960, for example, only six cities had land use regulations.

Figure 1 - Evolution of the number of municipalities with zoning ordinances.



Source: The authors, based on MUNIC 2013 data.

In addition, it can be observed that, from the 1980s on, there has been a great increase in the number of municipalities that have adopted zoning as an urban management instrument. Several interrelated factors can explain this growth. Firstly, it may be a natural consequence of the ongoing urbanization process in the country: in 1980, about 67.5% of the population already lived in urban areas. Moreover, in line with Gyourko, Saiz, and Summers (2008), the increase in wealth and education of Brazilian citizens can explain a greater awareness of the intensity and the use of the urban land. At that time, some institutional changes in the country were also associated with greater diffusion of zoning among municipalities. In 1979, the federal government took on a coordinating role of urban policy in cities. In this sense, the Federal law 6799/79 came into force, establishing general rules and guidelines for the division of urban land. This law also established the minimum lot size of 125 square meters, delegating to municipalities the option to adopt stricter criteria. Furthermore, in 1988, a new federal constitution was created, which, because of its decentralized characteristics, favored the autonomy of the municipalities in formulating their own public policies and the management of fiscal resources.

Due to intense city growth and to rural-urban movements, there was a strong increase in housing demand without a counterpart of the supply, causing a rise in the housing deficit[6] and the formation of slums across the country. To alleviate this problem, and realizing that the urban land restrictions can cause higher housing prices, some municipalities have adopted a kind of inclusionary zoning: the Special Zones of Social Interest (ZEIS)[7]. These are constituted of demarcated areas within the cities with more flexible and more specific rules regarding zoning. Other important institutional framework for the country's urban planning policy was the approval of the Federal Law 10257/01, known as "*Estatuto da Cidade*" which, among other things, emphasizes the social function of property, encourages participatory planning and enforces the adoption of a city master plan by a set of municipalities. This federal law may also have driven the strong adoption of zoning by the cities in the most recent period, from 2000 to 2013, where 1,013 new municipalities have adopted the order (figure 1).

Note that, although the zoning ordinances are present in 1,724 Brazilian municipalities (about 31% of the total), little is known about their impact on the formal housing market. In this work, we will take advantage of the variability of orders between the municipalities to verify the impact of land-use regulations on average rents and on the real estate market growth.

2.4. Empirical Strategy

2.4.1. Propensity Score Matching

The aim of this study is to investigate the effect of municipal zoning policies on average rent prices and on the real estate market growth. The great empirical challenge of this kind of analysis is to eliminate the possibility of endogeneity by reverse causality and by omitted variables. Although it is assumed that property prices can be explained by zoning ordinances, the reverse way is also possible: maybe zoning occurs just in areas that have more highly valued houses (Quigley, 2007). To reduce these concerns, we will apply a matching-based strategy in conjunction with a series of sensitivity analyses.

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[6] According to calculations of the João Pinheiro Foundation (2010), the housing deficit in Brazil is of 6,490 million housing units.

[7] 920 municipalities adopt the ZEIS through specific legislation, while 1,556 municipalities include the ZEIS as a part of the city master plan (MUNIC, 2013).

In this sense, our goal is to make a match between treated observations (municipalities that have implemented zoning ordinances) and the control observations[8] (municipalities that have not adopted zoning ordinances) such that the control group can be considered a counterfactual - what would happen to the treatment group had they not adopted zoning. Consider Z a binary variable that describes zoning. Thus, $Z = 1$ if the municipality has implemented the ordinance and $Z = 0$ otherwise. Y_0 and Y_1 describe the potential outcomes that vary according to the treatment (zoning order). In real life, only one of these outcomes can be actually observed for each municipality. However, the causal effect is defined as the difference between Y_1 and Y_0. We are interested in obtaining an estimate of the average effect of the treatment on the treated (ATT), defined as follows:

$$ATT = E(Y_1 - Y_0|Z = 1) \tag{1}$$

Under a set of assumptions, various methods can be used to estimate the ATT. One of these is the Propensity Score Matching (PSM) developed by Rosenbaum and Rubin (1983). The propensity score is the probability of receiving the treatment (in our case, implementing a zoning ordinance) given the observed covariates (defined as the X vector), i.e. $P(X) = P(Z = 1|X)$. The key identification assumption for the ATT is that Y_1 and Y_0 are independents of Z given $P(X)$. In other words, $Y_1, Y_0 \perp Z|P(X)$. This is known as Conditional Independence Assumption (CIA). Additionally, it is also necessary that, for each treated unit, a matched control unit with similar X exists. This assumption is known as overlap condition, and can be written as: $0 < P(Z = 1|X) < 1$. Finally, assuming that the CIA and the overlap condition holds, the impact of zoning on the city's average rental price which implemented the ordinance is given by:

$$\begin{aligned}\tau_{ATT} &= E(Y_1 - Y_0|Z = 1) = E[E(Y_1 - Y_0|P(X), Z = 1)] \\ &= E[E(Y_1|P(X), Z = 1) - E(Y_0|P(X), Z = 0)|Z = 1]\end{aligned} \tag{2}$$

That is, under these assumptions, we can use the outcomes of matched control municipalities as a counterfactual for the estimation of ATT. A two-step procedure is employed to obtain an empirical counterpart of (2). Firstly, a probabilistic regression model is estimated to get the propensity scores, $P(X)$. Second, based on algorithms, we make a match between the treated and non-treated municipalities which have a similar propensity score. For robustness purposes, we will apply three types of matching algorithms commonly used in the empirical literature: the nearest neighbor, the radius matching and the kernel matching. Caliendo and

[8] In Figure 1A of Appendix, we show the spatial distribution of municipalities by zoning adoption.

Kopeinig (2008) describe these algorithms in detail and discuss the advantages and disadvantages (in terms of bias and efficiency) of each one.

2.4.2. The plausibility of the CIA and sensitivity analysis

The CIA assumption implies that the adoption of zoning is completely explained by observable variables. That is, the matching procedure is constructed by using all covariates that generate differences in the distribution of treated and non-treated units, making the probability of treatment equal between the matched pairs, as if there had been a random assignment (Caliendo and Kopeinig, 2008). Thus, it is essential that the vector X contain detailed information on the relevant characteristics of municipalities, which will be possible with the Brazilian Demographic Census and the Survey of Basic Municipal Information (MUNIC). We will describe the variables used to obtain the propensity score in the next subsection. However, even with detailed information on the municipalities, there is a possibility that the CIA is not valid due to the existence of unobservable variables (confounders) that affect both the adoption of zoning and the outcome variable. The existence of confounders can make the ATT estimator biased and therefore hardly credible - this is the weakness of PSM strategies. Although the CIA is, by construction, a non-testable assumption and the absence of confounders is little convincing, we can - through a sensitivity analysis – either evaluate the strength that an omitted variable must have to invalidate our ATT estimate or simulate potential confounders. In this sense, we will use three different kinds of sensitivity analysis (Rosenbaum, 2002; Inchino et al., 2008 and Oster, 2016) to verify the robustness of the ATT when the Conditional Independence Assumption fails.

Firstly, we will apply the sensitivity analysis developed by Rosenbaum (2002), which provides an indication of the magnitude of the omitted variable bias that would be necessary to invalidate the associations initially observed by the estimated ATT. This method is based on the parameter Γ, which measures the difference in odds of receiving the treatment between observations with the same observable characteristics. Therefore, a random experiment ensures $\Gamma = 1$, and when that Γ grows, the experiment becomes more distant from randomization. For example, in an observational study, if $\Gamma = 3$, then one of the units has three times as likely to receive the treatment due to unobservable factors, since the treated and untreated units are identical regarding the observable variables. Thus, the basic procedure for Rosenbaum sensitivity analysis is to select a series of Γ values and, for each one, calculate a bound of significance level (p-values) for the ATT in the case of a endogenous treatment selection (Diprete and Gangl 2004). In addition, we will also apply a sensitivity analysis proposed by Ichino et al. (2008). This strategy aims to verify the bias of the estimated ATT when the CIA fails in some specific and relevant way. First, it is necessary to establish values for the parameters that characterize the distribution of a specific confounding factor, denoted by U. It's assumed that this variable is binary, independent and identically distributed. Thus, the distribution of U is completely characterized by the choice of four parameters, denoted by p_{ij}:

$$p_{ij} \equiv \Pr(U = 1|Z = i, I(Y > \bar{y}) = j) = \Pr(U = 1|X, Z = i, I(Y > \bar{y}) = j)$$
(3)

Where $i, j \in \{0,1\}$, I is an indication function and $\bar{y}$ is the mean of the outcome variable. Based on the parameters defined by (3), a value of U is given for each treated and control observation in accordance to one of the four groups to which the observation belongs. The simulated U is considered as a new observable covariate and is then included in the set of variables used for the matching process. Finally, the ATT is computed again with the inclusion of confounder U. Therefore, by modifying the parameters that characterize the distribution of U, we can evaluate the robustness of the ATT against different assumption related to the nature of confound factor (Nannicini, 2007). We will follow the simulation exercises proposed by Ichino et al. (2008) and choose p_{ij} in order to make the confounder distribution similar to the empirical distribution of important covariates. Furthermore, the confounders will be simulated to be threatening to the baseline ATT: will have positive correlation with the treatment status variable (Z) and with the outcome of non-treated units.

The third method we will apply is the sensitivity analysis recently developed by Oster (2016). Differently from the Rosenbaum (2002) and Ichino et al. (2008), the method is not grounded on PSM strategies and can be applied to OLS estimates with continuous treatment variables. Following the idea proposed by Altonji et al. (2005) that including an additional unobservable variable would have a similar effect in reducing selection bias as including an additional observable variable, Oster (2016) developed a consistent estimator of the treatment effect adjusted by the omitted variable bias, considering the assumption that there is a proportional selection between observable and unobservable variables:

$$\delta \frac{Cov\ (Z, X)}{Var\ (X)} = \frac{Cov\ (Z, U)}{Var\ (U)} \tag{4}$$

That is, the correlation between the observable variables (X) and the treatment variable (Z) is proportional to the correlation between the unobservable variables (U), and the treatment variable and the parameter δ is the coefficient of proportionality. Before implementing the estimator, Oster (2016) define the coefficient resulting from the simple regression of Y on Z as β' and the corresponding R-square as R'. Similarly, the coefficient from the intermediary regression of Y on Z and X is denoted by β'', while the R-square is given by R''. Finally, the R-square of a hypothetical regression of Y on Z, X and U is denoted as R^{max}. These parameters are used in the following relationship, which allows us to approximate the adjustment bias of the OLS regression:

$$\beta \approx \beta'' - \delta \frac{(\beta' - \beta'')(R^{max} - R'')}{(R'' - R')} \tag{5}$$

From the expression (5), it is possible to obtain an unbiased estimator for the treatment effect by calculating the bias given by the second term on the right side of the equation. Although δ and R^{max} are unknown parameters, Oster (2016) proposes some upper bounds. It is claimed that the selection in the unobservable should not exceed the selection in the observables, so that the maximum value of δ tends to be less or equal to one. In addition, Oster (2016) also shows that the value of $1.3R''$ is a suitable upper limit for R^{max}. To check the robustness in a practical way, we can construct a bounding set as $\Delta = [\beta'', \beta^*(R^{max}, \delta = 1)]$ and verify if it contains coefficients that move toward zero. Another way to check robustness is to find out the value of δ that would produce $\beta = 0$, corresponding to the degree of selection on unobservables relative to the observables that would be required to fully explain the treatment effect.

2.5. Variables and Data

As previously discussed, our empirical strategy will take advantage of the fact that some municipalities have adopted zoning ordinance while others do not. Information on urban policies that each city adopts are obtained from the survey of Basic Municipal Information (MUNIC), a database elaborated annually by the Brazilian Institute of Geography and Statistics (IBGE) to gather information on the dynamics, structure and functioning of local institutions. Thus, our treatment variable is dichotomous and takes the value of one if the municipality had a zoning law in 2004[9], and takes the value of zero otherwise. For robustness

9 We opted to choose this particular year because the information available in the MUNIC 2004 is consistent with each consulted local zoning order. Information about zoning are also collected in MUNIC 1999, 2005 and 2009. However, for these surveys there is some degree of disagreement with the local zoning orders. Anyway, in section 6.3, we will evaluate the robustness of the results to changes in the treatment assigned in the 1999, 2005 and 2009 MUNIC.

purposes, we also employ other treatment variable that measures the degree of rigidity of the minimum lot size endorsed by the municipality. It will take the value of one in the case of the local administration set a minimum lot size greater than 125 m² and zero otherwise.

Regarding the outcome variables, we will use two throughout the paper: the average rental prices (measured in 2010 and adjusted by the property quality) and the growth of formal houses in the 2000-2010 period. We will follow the Caliendo and Kopeing (2008) recommendation, and add only covariates that simultaneously affect the outcomes and the zoning adoption. To ensure that these variables are not directly affected by the treatment, they will be fixed in time or measured before the treatment (year 2000). The outcomes and the control variables were obtained from the Brazilian Demographic Census of 2000 and 2010. It is necessary to point out that since our outcome variable was measured only in 2010, it is impossible to build a panel database.

The choice of the control variables that will be used to proceed with the matching between the municipalities that adopt zoning and those that do not adopt was based on recent empirical literature (e.g. Ihlanfeldt 2007, Glaeser and Ward 2009; Zabel and Dalton 2011; Monkkonen and Ronconi 2013; Kok, Monkkonen and Quigley 2014). We split the covariates into five - not mutually exclusive - sets: I) Demand Factors: characteristics that drive the intense use of urban land, such as population density, population growth (1991-2000) and per capita income; II) Externalities Factors: activities that inflict damage to the natural and social environment of cities, causing greater pressure for the adoption of urban planning laws, which include urbanization and industrialization rates; III) Demographics: since the zoning orders are set by municipal legislators, local citizens have a strong influence on the conduction of urban policy. In this sense, we include the following variables: average years of schooling, the percentages of blacks, of working age, and immigrants, voter’s turnout, and percentage of homeowners; IV) Urban Infrastructure: as discussed in section 2, policies that restrict the urban land use can be simply a way to avoid high densities in places with poor urban infrastructure, thus we include the following variables in the control set: percentages of households with sewage, with electricity, and pipe water; V) Environmental and Fixed Factors: urban planning can also be established in order to protect the natural and cultural heritage from the real estate development. Thereby, cities that have coastline, units of natural conservation or historical-cultural capital may be more likely to adopt zoning orders. Therefore, we will use the year the city was founded as a proxy for the historical-cultural capital, coastline and conservation area dummies, and, finally, the distance to the nearest capital city. In addition to these sets of covariates, we will also include state-fixed effects in the propensity score estimation. The use of this broad set of variables that characterize the municipalities aims to make the distribution of treated and untreated municipalities as similar as possible. However, it is very likely that there are unobservables that affect the zoning assignment. Our argument in this study is that these variables are not important enough to invalidate our results.

An important concern in our analysis is the possibility that zoning occurs in cities with better housing standards. In this case, a positive relationship between zoning and average rental price could reflect only a simple association between prices and their hedonic attributes. In this study, we will use the correction proposed by Quigley and Raphael (2005), which seeks to adjust the rental prices in relation to property characteristics. This strategy aims to eliminate or to attenuate the influence of residences attributes of their respective rental prices, in order to obtain an average rent that only reflects the city's characteristics. Therefore, in order to obtain the outcome variable, the first step is to estimate a hedonic regression for each Brazilian municipality (5,507 separate regressions), where the dependent variable is the rental price and the independent variables are the properties characteristics available in the Demographic Census of 2010: number of rooms, number bathrooms, number of other rooms and a dummy that takes the value of one if the house is made of bricks and zero, otherwise (Seabra and Azzoni 2015). The dwelling is the observation unit in this analysis. In Appendix, the Table A1 summarizes the mean of the estimated coefficients. As expected, all the hedonic characteristics affect the rental prices positively. The last step is to combine the estimated coefficients of each city with the average hedonic characteristics of Brazil, in order to estimate the average rental price with constant quality. This will be our main outcome variable.

2.6. Results

In this section, we will present and discuss the results obtained by applying the methodology described above. First, we will evaluate the matching quality and estimate the ATT considering different algorithms (subsection 6.1); then, we will conduct sensitivity analyses (subsection 6.2); and, finally, we present some extensions and robustness checks (subsection 6.3).

2.6.1. Baseline Results

Table 1 shows the means of dependent variables and covariates by treatment status before and after the application of matching procedure. Initially, we can see that cities that adopt zoning order are substantially different from those that do not. For example, regulated cities have greater rental prices, higher house growth, higher per capita income, higher population density, are more urbanized, supply a better infrastructure, are closer to the state capital, and also have more natural amenities. It is noted that the mean difference for all observable characteristics is large and statistically significant between the two types of cities. This shows that a simple mean difference of rental price between municipalities with and without zoning is useless for impact evaluation.

The columns (4) and (5) of Table 1 show the variables' means after application of a nearest neighbor matching algorithm with replacement and imposing a common support. By applying this expedient, we found that 35 treated municipalities (0.63% of the total group) were outside of the common support region and, therefore, were excluded. We can see that after the matching, the characteristics of both kinds of municipalities become very similar, so that there are no statistically significant differences[10], suggesting a good matching quality.

Table 1 – Mean characteristics of cities that adopt zoning (treated) and that do not (control)

Variable Name	Before Matching			After Matching		
	(1)	(2)	(3)	(4)	(5)	(6)
	Treated	Control	Difference	Treated	Control	Difference
Log Average Rental Prices	5.652	5.220	0.432***	5.640	5.586	0.054**
House Growth	0.336	0.272	0.064***	0.343	0.353	-0.009
Log Income (Per Capita)	6.104	5.511	0.593***	6.086	6.070	0.015
Log Population Density	3.866	2.908	0.958***	3.777	3.732	0.045
Population Growth (91-00)	0.130	0.077	0.053***	0.128	0.116	0.012
% Workers in Manufacturing	0.213	0.137	0.077***	0.213	0.209	0.004
Urbanization Rate	0.715	0.546	0.169***	0.708	0.703	0.005
% Immigrants	0.436	0.338	0.098***	0.434	0.425	0.009
% Homeowners	0.752	0.774	-0.022***	0.752	0.756	-0.004
% Working-Age population	0.433	0.376	0.057***	0.432	0.433	0.000
% Black Population	0.047	0.062	-0.016***	0.047	0.048	-0.002
% Households with Pipe Water	0.691	0.544	0.147***	0.685	0.692	-0.007
% Households with Electricity	0.952	0.841	0.111***	0.951	0.954	-0.003
% Households with Sewage	0.302	0.204	0.099***	0.297	0.294	0.003
Avg. Schooling of 25 years old	5.121	3.673	1.448***	5.057	5.053	0.004
Voter Turnout in Mun. Elections	0.874	0.865	0.010***	0.875	0.877	-0.002
Year of Mun. Foundation	1957.4	1963.3	-5.900***	1957.8	1959	-1.200
Distance to the State Capital	233.520	259.860	-26.340***	237.400	245.360	-7.960

[10] We also use 27 state dummies to build the matching sample of Table 1. The results have been omitted for space constraints, but reveal that the mean difference (with respect to these variables) is not statistically significant between the two kinds of municipalities.

Coastal Municipality (0/1)	0.088	0.036	0.053***	0.079	0.096	-0.017
Conservational Area (0/1)	0.264	0.081	0.183***	0.249	0.244	0.005
Number of Observations	1386	4112	-	1351	713	-

Note: For the described matching, we used the nearest neighbor propensity score with replacement and imposing common support area. The average rent prices are corrected by their hedonic attributes (subsection 5.2). *** $p < 0.01$, ** $p < 0.05$, *** $p < 0.1$.

Table 2 summarizes the results of the propensity score estimation for unmatched and matched samples. We used a logit probability model with state fixed-effects. Although the choice of a probabilistic model is not critical to the ATT estimation, when compared to the probit function, the logit function has more density mass in the bounds (Caliendo and Kopeing 2008). In columns (1) and (2) we note that most of the estimated coefficients take the expected signs based on what has been discussed above and are consistent with the descriptive statistics of Table 1. Cities with higher income and more educated individuals are more likely to adopt zoning as an urban planning law, which conforms with Gyourko, Saiz, and Summers (2008). Areas that have a greater demand for land use (measured by population density), that are more urbanized, which have a higher manufacturing share, and, a precarious urban infrastructure also adopt zoning more easily. This last fact suggests that Brazilian cities may be using the zoning law as a legal instrument to limit development in areas with poor infrastructure. Additionally, we can observe a positive correlation between the use of zoning and the existence of natural amenities.

In column (3) and (4) of Table 4, the logit model is estimated considering only the sample of matched municipalities. The results show that the matching strategy performs well in simulating the randomness of the treatment variable, since the degree of explanation of the covariates on the zoning drops significantly, such that R^2 reaches to only 1.5%. At the bottom of Table 2, we also report other measures of matching quality suggested by Rosenbaum and Rubin (1983). Note that after the matching, the mean standardized bias dropped by about 87.7% (from 32.4 to 4). These results support those obtained in Table 1, indicating that, in fact, it is possible to build groups of reasonably similar municipalities.

Table 2 - Propensity Score Estimation: Determinants of Zoning

	Unmatched Sample		Matched Sample	
Dependent Variable: Zoning	(1)	(2)	(3)	(4)
	Coefficient	Std. Err.	Coefficient	Std. Err.
Log Income (Per Capita)	0.724***	0.219	0.344	0.244
Log Population Density	0.314***	0.053	0.074	0.052
Population Growth (1991-2000)	0.273	0.213	0.022	0.218
% Workers in Manufacturing	0.589	0.533	0.033	0.525
Urbanization Rate	0.656*	0.360	0.068	0.385

% Immigrants	-0.312	0.462	0.858	0.495
% Homeowners	-0.805	0.549	-0.518	0.623
% Working-Age population	-0.647	1.044	-1.003	1.094
% Black Population	0.036	1.197	-0.313	1.452
% Households with Piped Water	-0.128	0.299	-0.238	0.320
% Households with Electricity	-1.484***	0.634	-0.403	0.781
% Households with Sewer	-0.351	0.262	-0.186	0.257
Avg. Schooling of 25 years old	0.702***	0.091	-0.176	0.196
Voter Turnout in Mun. Elections	-2.778***	0.844	-0.508	0.900
Year of Mun. Foundation	-0.001	0.003	-0.007***	0.003
Distance to the State Capital	0.001**	0.000	0.000	0.000
Coastal Municipality (0/1)	0.585***	0.198	-0.343**	0.171
Conservational Area (0/1)	0.568***	0.114	0.069	0.100
Constant	-4.085	5.397	14.038***	5.406
State Fixed-Effects	Yes		Yes	
Pseudo R²	0.32		0.015	
LR chi2	1984.52***		32.44	
Mean Bias	32.4		4	
Median Bias	23.4		3.1	

Note: The dependent variable is an indicator for zoned municipalities. The estimation was carried out with a logit function. *** p< 0.01, ** p < 0.05, *** p < 0.1.

After obtaining the propensity scores, we calculate the impact of zoning on the average rental prices and the real estate market growth using four different matching algorithms. For rental, Table 3, Panel A, shows the estimated ATT and the corresponding measures of matching quality described above. The estimated ATTs vary from 0.054 to 0.062, indicating that cities that adopt zoning as urban planning policy have an average rent 5% higher compared to the non-adopters. This effect is substantial, since it corresponds to about 12.46% of the unconditional difference of average rental between the two kinds of municipalities. Based on our discussion of Section 2, this evidence shows that the zoning orders drawn up by Brazilian municipalities have some level of compliance. In addition, the positive impact of zoning on prices is also in line with the empirical literature that, when analyzing American cities, found positive associations between higher levels of land-use restrictiveness and real estate prices. (Quigley and Raphael 2005, Ihlandfeldt 2007, Zabel and Dalton 2011).

Table 3 – The Impact of Zoning on Average Rental Prices and on Real Estate Market Growth

Panel A: Average Rent Prices	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0539**	0.0584***	0.0601**	0.0626***
	(0.0212)	(0.0191)	(0.0218)	(0.0201)
	[0.0164]	[0.0110]	[0.0113]	[0.0098]
Mean Bias After Matching	3.1	1.5	1.8	1.6
Pseudo R²	0.015	0.009	0.008	0.008

Panel B: House Growth	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	-0.0094	0.0071	0.0012	0.0079
	(0.0393)	(0.0280)	(0.0265)	(0.0255)
	[0.0356]	[0.0297]	[0.0284]	[0.0253]
Mean Bias After Matching	3.7	2.4	2.7	2.6
Pseudo R^2	0.019	0.009	0.01	0.008

Note: Analytical and bootstrap standard errors (200 interactions) are reported in round and square brackets, respectively. The Radius matching uses a 0.01 caliper and the Kernel matching is constructed using an Epanechnikov function. *** $p < 0.01$, ** $p < 0.05$, *** $p < 0.1$. The average rent prices are corrected by their hedonic attributes (subsection 5.2).

The panel B of Table 3 shows that, regardless of the matching algorithm used, the zoning law does not affect real estate market growth. Therefore, we cannot associate the higher rental prices simply to a reduction in housing supply. There is a variety of mechanisms that can explain the causes of larger rental prices in regulated cities. For example, the need to comply with parameters and building codes can generate higher development costs (Quigley 2007), higher land prices (Kok, Monkkonen and Quigley 2014), or the land-use regulation can reduce housing supply elasticity, which favors large increases of prices in response to variations in demand. In addition, zoning can function as a kind of insurance against conflicting land uses, thus causing an appreciation of proprieties (Zhou, McMillen and McDonald 2008). While understanding the mechanisms that underlie the relationship between zoning and higher prices is important, the aggregated nature of our database does not allow inferences about them.

In order to investigate how reliable our estimates are, we consider how the estimated ATTs change assuming failures in Common Independence Assumption (CIA). We believe that, even including important covariates to the propensity score estimation, there are potentially immeasurable variables that affect both the outcome variable and zoning adoption. For example, the degree of concern of municipal governments with issues related to urban planning is not measurable and is a potential confounding factor that can invalidate the CIA. In the next subsection, we will conduct different kinds of sensitivity analysis.

2.6.2. Sensitivity Analyses

Firstly, we employ the Rosenbaum (2002) bounds approach to measure the sensitivity of the baseline specification when our key assumption (CIA) is relaxed. Thus, we use different values of Γ -which measures the difference in odds of receiving the treatment between observations with the same observable characteristics - to verify changes in inference due to the existence of unobservable confounders. Table 4 shows the results for Γ ranging from 1 to 2 and the corresponding p-value bounds.

It can be seen that up to $\Gamma = 1.8$, the estimated ATT remains statistically significant at the usual level of 5%. This indicates that a city may have a chance to adopt zoning 80% higher than the others due to unobservable characteristics and yet, our initial evidence remains reliable. Only when $\Gamma = 1.9$, the ATT is no longer significant at the usual levels. Thus, for the higher rental prices to be associated with a unobservable variable instead of zoning order it is necessary that this confound increase by 90% the odds of a municipality implement the zoning. Thus, the results of Table 3 cannot be explained by a simple moderate omitted-variable bias.

Table 4 - Sensitivity Analysis I: Rosenbaum Bounds

Average Rent Prices					
Γ	p-crit+	p-crit-	Γ	p-crit+	p-crit-
1.00	0.000	0.000	1.55	0.000	0.000
1.05	0.000	0.000	1.60	0.000	0.000
1.10	0.000	0.000	1.65	0.001	0.000
1.15	0.000	0.000	1.70	0.004	0.000
1.20	0.000	0.000	1.75	0.014	0.000
1.25	0.000	0.000	1.80	0.039	0.000
1.30	0.000	0.000	1.85	0.088	0.000
1.35	0.000	0.000	1.90	0.170	0.000
1.40	0.000	0.000	1.95	0.285	0.000
1.45	0.000	0.000	2.00	0.426	0.000
1.50	0.000	0.000			

Note: the parameter Γ measures the odds of receiving the treatment considering municipalities with the same observable characteristics. The method is built based on the nearest neighbor matching without replacement.

The results of the sensitivity analysis proposed by Inchino et al. (2008) can be seen in Table 5. We choose values of p_{ij} in order to simulate confounders with distributions analogous to those of the observed covariates and that have a positive correlation with treatment and outcome variables[11]. This way, we simulated unobservables with a similar distribution of the following variables: per capita income, population density, urbanization rate, average schooling, conservation area and coastal municipality dummy. It is worth mentioning that the Inchino et al. (2008) procedure assumes that the distribution of simulated confounder is discrete[12]. Thus, to mimic the continuous covariates, we use the following

[11] This would be an unobserved variable with potential to explain the positive impact of zoning on rental prices.

[12] Through Monte Carlo simulations, the authors concluded that the assumption that the confounder assumes a discrete distribution when in fact it is continuous is not able to generate erroneous conclusions regarding the sensibility of the ATT. Moreover, Wang & Krieger (2006) show that binary variables are more threatening to causal inference than continuous ones.

criteria: the simulated binary variable will take the value of one if its value is above the mean (or median) of the distribution, and zero otherwise.

Table 5 - Sensitivity Analysis II: Simulated Confounders

	p11	p10	p01	p00	Out. Eff.	Selec. Eff	ATT	SE
Baseline	0	0	0	0	-	-	0.054***	0.021
Neutral	0.5	0.5	0.5	0.5	1.019	1.006	0.062***	0.008
Confound to Mimic								
Income (Mean)	0.94	0.44	0.76	0.16	17.598	3.713	0.044***	0.010
Income (Median)	0.91	0.35	0.71	0.13	17.736	3.406	0.045***	0.009
Average Schooling (Mean)	0.92	0.43	0.71	0.14	15.146	4.044	0.044***	0.010
Average Schooling (Median)	0.92	0.42	0.69	0.13	15.338	4.075	0.043***	0.010
Density (Mean)	0.71	0.49	0.45	0.45	0.993	2.327	0.062***	0.010
Density (Median)	0.7	0.49	0.44	0.45	0.990	2.300	0.061***	0.010
Urbanization Rate (Mean)	0.77	0.46	0.59	0.32	3.111	2.193	0.055***	0.010
Urbanization Rate (Median)	0.76	0.45	0.58	0.31	3.120	2.190	0.054***	0.009
Coastal City (1/0)	0.09	0.08	0.04	0.03	1.286	2.545	0.062***	0.008
Conservational Area (1/0)	0.29	0.11	0.12	0.05	2.945	2.993	0.058***	0.011

Note: The method is based on nearest neighbor matching with 200 interactions. The outcome effect measures the impact of the unobserved variable on the untreated outcome, while the selection effect measures the impact of the unobserved variable on treatment assignment. *** $p < 0.01$, ** $p < 0.05$, *** $p < 0.1$.

From Table 5, it can be seen that when the CIA does not hold due to the existence of a specific unobservable confounder, the magnitude of the ATT changes but remains qualitatively important and statistically significant. For example, even if we are omitting an unobservable that is too important to the adoption of zoning such as per capita income or population density (Alterman 2013, Gyourko, Saiz and Summers 2008), our initial evidence remains valid.

Finally, Table 6 presents the results of the Oster (2016) procedure for different values of R^{max}. We estimate two types of OLS specifications to generate movements in the R-Squared in order to calculate the selection bias given by equation (5): the partially controlled model, which includes only the treatment variable and the municipality average schooling and the fully controlled model, which includes the full set of controls. The first line of Table 6 shows the bounding set considering equal selection between observable and unobservable variables ($\delta = 1$) and the second line shows the δ values that would be necessary to invalidate the conclusion that zoning affects the rent prices.

Table 6 - Sensitivity Analysis III: Oster's Procedure

	$R^{max} = 0.8$	$R^{max} = 0.85$	$R^{max} = 0.9$	$R^{max} = 0.95$	$R^{max} = 1$
Bounding Set with $\delta = 1$	[0.049, 0.054]	[0.044, 0.054]	[0.039, 0.054]	[0.034, 0.054]	[0.028, 0.054]
Value of δ for $\beta = 0$	6.775	3.866	2.705	2.08	1.69

Note: The method is based on OLS estimates. For the partially controlled model, $\beta' = 0.068$ and $R' = 0.54$ and for fully controlled model, $\beta'' = 0.054$ and $R'' = 0.735$.

Firstly, we noted that for any maximum value of R-square, the bounding set does not include zero. Therefore, the ATT remains valid in all situations assuming that the unobservables confounders are as important for the selection as the observables. Moreover, even considering the threshold of $R^{max} = 1.3R''$ suggested by Oster (2016), which is equivalent to 0.95, it would be necessary that the selection in unobservables was 2.08 higher than the selection in the observables for invalidate our conclusions. These set of evidences reinforce the conclusion reached by applying the Rosenbaum bounds and the simulated confounders, indicating that, in fact, the positive impact of zoning on the average rental price does not seem to be driven by some omitted variable.

2.6.3. Robustness Checks

In addition to the sensitivity analysis described in the previous subsection, we also verified the robustness against changes in the definition of treatment and outcome variables and the robustness against restricted samples of municipalities. Firstly, it was considered a more specific measure of land-use restrictiveness: the minimum lot size. As discussed in Section 2, some municipalities determine a minimum lot size greater than the recommendation of Federal Law 6799/79, becoming more rigid. Thus, we modify the treatment variable to a dummy that takes the value of 1 in case the municipality set a minimum lot size higher than 125m² and 0 otherwise. To get a clearer impact of this regulatory instrument, we dropped the municipalities that already have some zoning ordinance. Thus, the number of municipalities fell to 4,121, where 956 (23.19%) set a minimum lot size higher than 125m². The results of this exercise can be seen in panel A of Table 7.

Table 7 – The impact of Zoning: ATT using different treatment and outcome variables

Panel A - Different treatment:	(1)	(2)	(3)	(4)

Higher Minimum Lot Size				
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0381**	0.0301***	0.0284***	0.0351***
	(0.0211)	(0.0169)	(0.0179)	(0.0173)
	[0.0177]	[0.0113]	[0.0107]	[0.0102]
Mean Bias After Matching	2.6	1	0.9	1.2
Pseudo R²	0.014	0.002	0.002	0.002
Panel B - Different outcome variable: Average Rent Prices	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.04607***	0.05125***	0.05112***	0.05462***
	(0.0227)	(0.0198)	(0.0222)	(0.0205)
	[0.0165]	[0.0111]	[0.0105]	[0.0088]
Mean Bias After Matching	3	2.3	2.5	2.4
Pseudo R²	0.013	0.008	0.008	0.008
Panel C - Different outcome variable: Median Rent Prices	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0295*	0.0398***	0.0407***	0.0442***
	(0.0242)	(0.0207)	(0.0229)	(0.0212)
	[0.0173]	[0.0107]	[0.0096]	[0.0092]
Mean Bias After Matching	3	2.3	2.5	2.4
Pseudo R²	0.013	0.008	0.008	0.008

Note: In panel A, the treatment variable is modified to a dummy that takes the value of 1 if the municipality adopts a minimum lot size (MLS) greater than 125 square meters and 0 otherwise. In panels B and C, the outcome variable is modified to the average/median rent prices without the hedonic correction. Analytical and bootstrap standard errors (200 interactions) are reported in round and square brackets, respectively. The Radius matching uses a 0.01 caliper and the Kernel matching is constructed using an Epanechnikov function. *** p < 0.01, ** p < 0.05, *** p < 0.1.

The estimated ATT indicates that in municipalities that adopt stricter criteria regarding the minimum lot size, the average rental prices are around 2.84% to 3.51% higher when compared to those who do not. This result is in line with our previous discussion and reveals the robustness of the results to different treatment criteria. In panels B and C of Table 7, we change the outcome variable to the average and median[13] rental price without making any adjustments to the property characteristics. That is, the rental prices are not measured with constant quality. This robustness test is useful to demonstrate that the adjustment made in the outcome variable, before calculating the zoning impact, was not decisive to find an economically significantly result. In this case, the Figures of panels B and C indicate that there was a reduction in the magnitude of the estimated ATT. However, the results remain economically important, indicating robustness.

Second, as argued in subsection 5.1, we chose to use the treatment assignment (zoning) based on the MUNIC survey of 2004. Data about zoning ordinances were also

[13] Unlike the average rent, the median rent is less sensitive to extreme values.

collected in MUNIC 1999, 2005 and 2009. However, among these surveys, there are some divergences regarding the information about the municipality zoning adoption. The 2004 survey has a higher degree of compliance with local laws, inspected manually. For this reason, this survey was chosen to define the municipalities that have zoning ordinances. Anyway, to demonstrate that our result was not driven by this choice, we also estimate the ATT modifying the treatment criterion to that one's defined on MUNIC 1999, 2005 and 2009. The results are displayed in table 2A of Appendix (Panels A, B and C). In panel D of Table 2A, the assignment is defined as follows: the municipality is treated (has zoning ordinances) if defined in at least one of MUNIC surveys. The result of table 2A shows that although there is a significant variation in the estimated ATTs, the results remain significant and economically relevant. That is, our conclusion is not sensitive to different MUNIC surveys.

Our last robustness test restricts the sample of municipalities considered in the previous estimates, since there is a possibility that a set of cities with urban structures very different from the Brazilian standard completely drive our results. For example, perhaps the positive impact of zoning is capturing only the unobservable differences between municipalities that are part of metropolitan regions (RMs) and rural municipalities. In the RM municipalities, there is a greater variety of economic sectors, which can lead to a differentiated dispute over land use and different claims for urban planning instruments. In addition, the municipalities of the central-west region and the northern region of Brazil are located in areas of more recent occupation, with a greater presence of environmental assets and lower urban density, which also can affect the results.

In order to check the robustness of our estimates in face of these divergences in urban structures, Table 8 shows the results of the ATT by removing the municipalities that are part of the metropolitan regions (Panel A), the municipalities of the state of São Paulo and Rio (they had a faster and deeper urbanization in relation to the rest of the country, Panel B), those of the Central West region (Panel C) and the municipalities of the North (Panel D). Panel E and F of Table 8 exclude from the sample cities that implemented inclusive zoning (known as ZEIS, as discussed in section 3 above) and municipalities with a large proportion[14] of irregular houses. In both cases, the real estate market of these two types of cities can potentially behave quite differently from the rest of the country, since there is a possibility of non-compliance with traditional zoning ordinances. The figures in Table 8 indicate that the

[14] We define municipalities with a large proportion of irregular houses as those with 30% or more of houses classified by IBGE as subnormal's.

results obtained previously are not significantly affected by the existence of municipalities with distinct patterns of urbanization, signalizing once more robustness.

Table 8 - The impact of Zoning: ATT using different samples of Municipalities

Panel A - Dropping RM Municipalities	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0528**	0.0610***	0.0566***	0.0671***
	(0.0215)	(0.0183)	(0.0194)	(0.0191)
	[0.0183]	[0.0108]	[0.0108]	[0.0108]
Mean Bias After Matching	1.8	1.2	1	1.2
Pseudo R^2	0.012	0.006	0.004	0.006
Panel B - Dropping SP and RJ Municipalities	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0634***	0.0619***	0.0594***	0.0669***
	(0.0248)	(0.0206)	(0.0234)	(0.0215)
	[0.0182]	[0.0124]	[0.0113]	[0.0113]
Mean Bias After Matching	4.8	1.5	2.3	1.4
Pseudo R^2	0.024	0.01	0.01	0.009
Panel C - Dropping Midwest Municipalities	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0623***	0.0560***	0.0589***	0.0619***
	(0.0241)	(0.020)	(0.0238)	(0.0214)
	[0.0157]	[0.0117]	[0.0105]	[0.0102]
Mean Bias After Matching	1.9	1.4	1.3	1.4
Pseudo R^2	0.011	0.008	0.007	0.007
Panel D - Dropping North Municipalities	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.04785**	0.04926***	0.0551***	0.05952***
	(0.0221)	(0.0200)	(0.0221)	(0.0208)
	[0.0165]	[0.0101]	[0.0107]	[0.0102]
Bias After Matching	1.9	1.5	2.2	2
Pseudo R^2	0.011	0.009	0.01	0.009
Panel E - Dropping ZEIS Municipalities	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0474**	0.0523***	0.0565***	0.0631***
	(0.0217)	(0.0186)	(0.0207)	(0.0198)
	[0.0165]	[0.0123]	[0.011]	[0.0104]
Bias After Matching	3.0	1.2	1.4	1.6
Pseudo R^2	0.018	0.005	0.006	0.006
Panel F - Dropping Municipalities with High Share of Informal Residences	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0535***	0.0537***	0.0590***	0.0614***
	(0.0209)	(0.0186)	(0.0219)	(0.0198)
	[0.0170]	[0.0111]	[0.0104]	[0.0087]
Bias After Matching	2.1	1.6	1.5	1.6
Pseudo R^2	0.017	0.009	0.009	0.008

Note: Analytical and bootstrap standard errors (200 interactions) are reported in round and square brackets, respectively. The

Radius matching uses a 0.01 caliper and the Kernel matching is constructed using an Epanechnikov function. *** $p < 0.01$, ** $p < 0.05$, *** $p < 0.1$. Average rent prices were corrected for their hedonic attributes (subsection 5.2).

2.7. Conclusions

In recent decades, due to the increased autonomy of local governments, the rising of income and a greater federal concern with urban planning issues, there has been, in Brazil, the diffusion of the idea of the urban land-use regulations. In this context, the number of cities with zoning law has grown from 64 in 1978 to 1,724 in 2013. Nevertheless, little is known about the impact of these local policies on the functioning of the Brazilian real estate market. Our study sought to contribute to this debate.

To reduce concerns about the endogeneity in the relationship between zoning and prices, we use a strategy based on the Propensity Score Matching (PSM). However, a necessary assumption is that the adoption of zoning is determined only by observable characteristics. This hypothesis, known as Conditional Independence Assumption (CIA), is little credible. Thus, through sensitivity analysis (Rosenbaum 2002, Inchino et al. 2008 and Oster 2016), we evaluate how our findings are robust to failures in the CIA. The results indicate that zoning generates an increase in the average rental price by 5.4% to 6.3%. Moreover, this evidence was robust to different matching algorithms, to the fails in CIA and to changes in the definition of the outcome and treatment variables.

The fact that the cost of housing is higher in cities that restrict land use is consistent with previous studies showing that regulations favored the development of the informal housing market and slum formation (Biderman, 2007, Cavalcanti and Da Matta 2013, Dantas et al. 2018). Facing a higher price in the formal housing market, low-income households tend to migrate to the informal ones. Consequently, there is a welfare loss, since this type of market is characterized by poor urban infrastructures. From this point of view, in the case of Brazilian municipalities, zoning ordinances can be generating social costs that are not negligible. Thus, under public policies perspective, Brazilian municipal land use restrictions can negatively affect the more recent social housing financing program Minha Casa Minha Vida, once its positive influence on prices turns the residential location financed by the federal program more distant from jobs.

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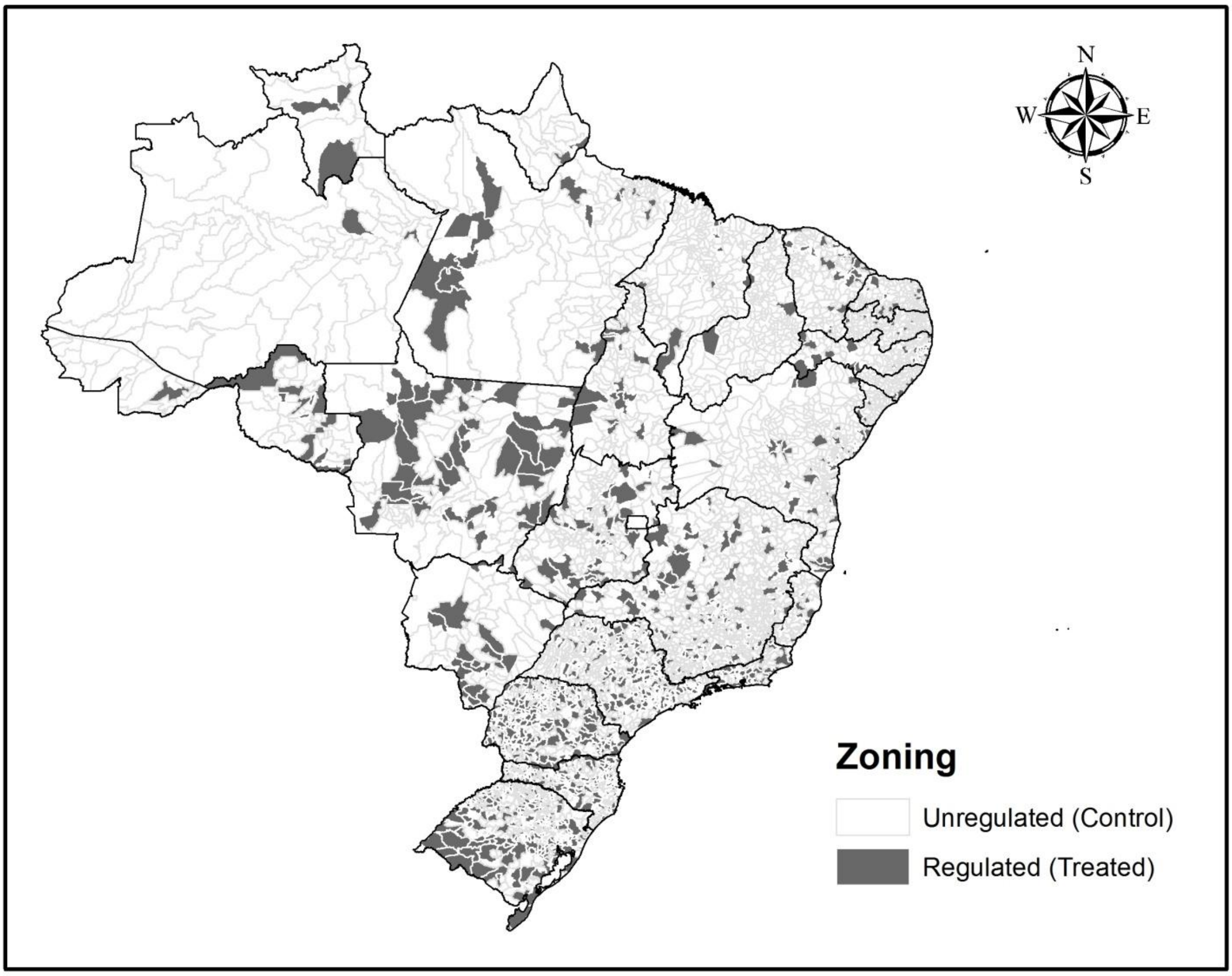
2.9. Appendix A – Second Chapter

Figure 1A – Spatial Distribution of Municipalities by Zoning Adoption



Source: Survey of basic municipal information (MUNIC), 2004.

Table 1A – Average Regression Coefficients of the Hedonic Models.

Variable	Average Coefficient
Brick Dummy (1/0)	74.4631
Number of Rooms	30.9680
Number of Bathrooms	218.7929
Number of Other Rooms	30.0606
Constant	-128.0507

Note: The figures reported in the table display the average coefficient for each variable. 5507 municipality-regressions were estimated, where the observation unit is the dwelling. The number of observations varies for each municipality.

Table 2A – The impact of Zoning: ATT using different MUNIC surveys to establish the treatment assignment.

A - Zoning MUNIC 1999	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.02053	0.03610***	0.03461***	0.05200***
	(0.0236)	(0.0203)	(0.0215)	(0.0222)
	[0.0175]	[0.0133]	[0.0105]	[0.0112]
N Treated	1186	1186	1186	1186
Off Support	34	34	72	34
Bias After Matching	3.6	2.7	2.4	2.8
Pseudo R^2	0.017	0.008	0.007	0.007
B - Zoning MUNIC 2005	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0481**	0.0342***	0.0399***	0.0462***
	(0.0214)	(0.0164)	(0.0166)	(0.0169)
	[0.0191]	[0.0106]	[0.0096]	[0.0097]
N Treated	1131	1131	1131	1131
Off Support	5	5	26	5
Bias After Matching	2.5	1.9	1.3	1.4
Pseudo R^2	0.010	0.004	0.002	0.003
C - Zoning MUNIC 2009	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0344***	0.0395***	0.0382***	0.0428***
	(0.0201)	(0.0161)	(0.0158)	(0.0157)
	[0.0130]	[0.0094]	(0.0084)	[0.0080]
N Treated	2111	2111	2111	2111
Off Support	3	3	30	3
Bias After Matching	3.2	1.9	1.7	2
Pseudo R^2	0.012	0.004	0.004	0.005
D - Zoning Sum	(1)	(2)	(3)	(4)
	Nearest Neighbor	10 Nearest Neighbor	Radius	Kernel
ATT	0.0421**	0.0637***	0.0488***	0.0706***
	(0.0225)	(0.0195)	(0.0218)	(0.0183)
	[0.0161]	[0.0112]	[0.0116]	[0.0094]
N Treated	2611	2611	2611	2611
Off Support	132	132	132	132
Bias After Matching	4.90	3.1	4.3	3
Pseudo R^2	0.02	0.012	0.017	0.01

Note: Analytical and bootstrap standard errors (200 interactions) are reported in round and square brackets, respectively. The Radius matching uses a 0.01 caliper and the Kernel matching is constructed using an Epanechnikov function. *** $p < 0.01$, ** $p < 0.05$, *** $p < 0.1$. Average rent prices were corrected for their hedonic attributes (subsection 5.2).

Chapter 3 - Secession of Municipalities and Economies of Scale: Evidence from Brazil[15]

3.1. Introduction

After the re-democratization process and the creation of a new constitution in 1988, Brazil experienced an intense process of political and administrative decentralization. According to the new constitution, municipalities formally became entities of the federation, and thus gained a greater degree of autonomy, which included the administration of resources accrued from local taxes and federal transfers (Gomes & Mac Dowell, 2000; Tomio, 2002a). Under this new institutional context, there was a continuous process of district[16] emancipations, with the number of municipalities increasing from 4,491 in 1991 to 5,561 in 2001.

It is interesting to note that this process of municipality secession in Brazil contrasts with the experiences of a significant number of developed countries that have reformed their municipal boundaries in order to amalgamate municipalities. For example, Sweden had 2,498 municipalities in 1951,but after some reforms this number shrank to 298 in 1974 (Jordahl & Liang, 2009, Hinnerich, 2009). In Israel, the number of municipalities decreased from 264 to 253 due to the 2003 amalgamation plan (Reingewertz, 2012). The list of countries that have adopted similar reforms includes the Netherlands (Allers & Geertsema, 2016), Finland (Saarimaa & Tukiainen, 2015), Germany (Blesse & Baskaran, 2016), Japan (Nakazawa, 2013), New Zealand (Kortt, Dollery, & Drew, 2015), Australia (Dollery, Byrnes, & Crase, 2007), Denmark (Blom-Hansen, 2010), and Canada (Vojnovic, 2000).

The main motivation in developing amalgamation plans is the gains from economies of scale, i.e., it is expected that public per capita expenditures will fall when the municipality size increases (Reingewertz, 2012; Fox and Gurley, 2006; Vojnovic, 2000). Some authors also point to other advantages of large cities, such as gains associated with political bargaining at higher levels of government, the possibility of economies of scope, the increase of the technical capacity of public agents and administrators, the possibility of reducing regional disparities and the ability to provide better regional planning (Fox & Gurley, 2006;

[15] Published on *Journal of Regional Science, 58*(1), 159-180.

[16] In Brazil, the districts are subdivisions of the municipalities.

Dollery, Byrnes & Crase, 2007). Additionally, a merger of municipalities may also increase the public-service level and the residents' welfare because it potentially reduces the tax competition between the localities. That is, as the number of jurisdictions falls, the negative externality associated with wasteful tax competition can be internalized (Hoyt, 1991). However, there are also arguments opposed to the merging of municipalities and in favor of the increased fragmentation of local governments. First, gains from large-scale economies are not valid for all services and public goods (Ladd, 1994; Fox & Gurley, 2006). Furthermore, proponents of decentralization argue that merging subnational governments has the potential of reducing competition between local governments and of generating rent-seeker monopolies – with high taxes and low efficiency (Nelson, 1992). Finally, Swianiewicz (2010) and Lassen and Serritzlew (2011) pointed out that in a set of smaller and more fragmented municipalities, the link between voters and their representatives is closer, making the degree of democracy and political accountability potentially higher.

The empirical evidence about the consequences of changing the number of subnational administrative unities is basically focused on amalgamation of the municipalities of developed counties, and points in mixed directions. Some papers find that it is an effective strategy for reducing per capita public expenditures through economies of scale. For example, Reingewertz (2012) analyzed the Israeli amalgamation reform of 2003 and showed that there was a 9 percent reduction in the expenditures of the municipalities that combined, and in Germany, Blesse and Baskaran (2016) detected a decrease of around 8.8 percent in the per capita administrative spending of the local governments that had merged. On the other hand, some papers point to increased spending or even no effect after amalgamation. Welling-Hansen (2011), analyzing the case of Denmark, and Moisio and Uusitalo (2013), investigating Finland's experience, for example, both found increases in public spending. And more recently, Allers and Geertsema (2016) found no significant effect on aggregate spending or taxation in the Netherlands.

Note that this empirical literature, in addition to focusing mainly on amalgamation experiences, does not generally consider the experiences of developing countries with local boundary reforms, where achieving higher efficiency in administrating public resources can be even more important, given the general low quality of public services. Particularly for the Brazilian experience, although the debate about the consequences of municipal secessions is quite intense internally, there are few studies that have investigated the issue. Gomes and MacDowell (2000) initially showed that this phenomenon could generate a greater transfer of resources from the large municipalities to the small ones (which are not necessarily the poorest), fostering economic distortions. More recently, Mattos and Ponczek (2013) found that there was deterioration in the supply of public goods and other social indicators in those municipalities that were sectioned, and Boueri et al. (2013) found a reduction of economic dynamism in their governments. The little available evidence thus appears to be unfavorable to the creation of new municipalities in Brazil.

This available set of evidence represents only indirect evidence about the potential impact of the secession process in Brazil, and in particular, it does not uncover the impact of municipal dissolutions on public expenditures or their relation to economies of scale. This study aims to fill this gap. More specifically, considering the debate about the economic implications of the amalgamation of municipalities, we investigate if the Brazilian secession of municipalities in the 1990's brought an increase in current and capital expenditures to the municipalities involved, and in what sense this potential increase can be associated with reduced economies of scale in the provision of public goods and services.

In order to investigate the causal impacts of the changed municipal boundaries, we use a Difference-in-Difference identification strategy, fixing the municipal boundaries at their pre-reform limits. The results, obtained using official data, indicate that those municipalities that underwent secession increased the local per capita expenditures on capital by 14.7 percent and per capita current expenditures by 7.8 percent. These results are robust to the consideration of different control groups and forms of model misspecifications. Furthermore, additional evidence indicates that increasing the per capita capital expenditures can hardly be explained by other channels than losing economies of scale (as opposed to availability of public goods, increased local revenues or attractiveness of municipalities) and rent-seeking gains for local leaders.

Besides this introduction, the paper is organized into five more sections. In the next section, we present the institutional background against which the emancipation of some municipal districts occurred. In section 3 we describe the identification strategy and methodological aspects of the work, and in section 4 we present the data and descriptive statistics. In section 5 we present and discuss the results, proceed to robustness checks and explore the possible mechanisms behind the secession of municipalities in Brazil. Final considerations are presented in section 6.

3.2. Institutional Background

After 20 years of military dictatorship in Brazil, a new federal constitution was established in 1988. In addition to its remarkable democratic aspects, administrative decentralization can be considered as one of the most important features of the new Brazilian constitution. It made the municipalities to be formal administrative entities, implying greater tax autonomy, decision-making freedom and the possibility of receiving a greater proportion of the public resources. In this context of increasing political decentralization, Article 18 of the Federal Constitution gives autonomy to the states to establish their own rules (through their legislative chambers) regarding the creation, merging and dividing of their respective municipalities.

The new constitution of 1988 specifically defined sources of income, and social responsibilities and competencies for the union, states and municipalities. While defining broader administrative competencies for the union and states, those for the municipalities from 1988 forward were associated with the provision of local services, including public transport, infant and basic education, health services, regulation of land use, and the conservation of history and culture (Tristão, 2003). On the other hand, municipal sources of income are composed of three kinds of receipts: local taxes, and involuntary and voluntary transfers of tax money from the union and states. Local taxes are levied on urban property (IPTU, *Imposto sobre a Propriedade Predial Urbana*), transfers of goods and property (ITBI, *Imposto sobre as Transmissões de Bens e Imóveis*), and services (ISS, *Imposto sobre Serviços*).

As pointed out by Afonso and Araújo (2000), however, based on official data considering all municipalities between 1988 and 2000, local taxes never reached more than 5.1 percent of the country's total tax revenue and never constituted the major part of total municipal tax income (which was around 17 percent of national tax income in 2000). In fact, even with the sizeable growth during the period, local municipal taxes were only 30.1 percent of total municipal taxes received in 2000. Thus, during this period, the greater part of the revenues received by municipalities came from transfers of tax money collected by the union and states, including, for example, the FPM (*Fundo de Participação dos Municípios*). Instituted by the 1988 constitution in order to transfer resources to the municipalities, it constitutes the main source of support for small municipalities.

This illustrates a characteristic of the fiscal decentralization in Brazil: the decentralization of competencies associated with the provision of local services was quicker than the local capacity of municipalities to collect taxes. Note, in addition, that because municipal taxes (mainly ISS and IPTU) tend to provide more benefit to larger and economically stronger municipalities, this imbalance between competencies and local capacity to collect taxes tends to vary among municipalities. By 1999, for example, the municipalities of the Southeast (the richest area) and Northeast (the poorest) collected, respectively, 71 percent and 8 percent of total Brazilian municipal own-receipts from taxes; in the same year, these shares were 55 percent and 19 percent, respectively, when considering total municipal receipts (Afonso and Araújo, 2000). This shows that federal and state fund transfers, including the more explicit FPM, play an important role in attenuating disparities among municipal receipts in Brazil.

Faced with these new fiscal and competency regulations and driven by other motivations (which will be discussed below), a large number of districts filed in their legislative assemblies to become new municipalities. Simply speaking, these processes consist of three distinct stages[17]: the passing of a plebiscite by the population concerned, a consideration and vote by the state legislature and finally the approval or veto of the executive power. Tomio (2002a) analyzed the political mechanisms – interactions between the legislature and governor– that described the rhythm of the emancipations in the ‘90s. The author argued that both the local political forces and legislatures had strong incentives to support the emancipation requests, so the only resistance would be from a contrary governor. As this situation was unlikely, the approval of emancipations was usually easy. Thus, there was a sharp increase in the number of Brazilian cities: one of about 24 percent in just 10 years (Table 1). In fact, within this period, we can highlight two moments of intense expansion: in 1993, when 483 new municipalities were established, and in 1997, when 533 new administrations were created[18].

Table 1 - Evolution in the number and size of Brazilian municipalities (1991-2010)

	1991	1993	1997	2001	2010	Variation
Brazil	4,491	4,974	5,507	5,561	5,566	24%
North	298	398	449	449	449	51%
Northeast	1,509	1,558	1,787	1,792	1,794	19%
Southeast	1,432	1,533	1,666	1,668	1,668	16%
South	873	1,058	1,159	1,189	1,189	36%
Midwest	379	427	446	463	466	23%
Average Population	32,693	30,473	28,988	30,164	34,278	5%

[17] Legislation on the establishment, merging and dismemberment of municipalities is the responsibility of the states, while the requirements for emancipation (such as population, voters and physical infrastructure) and lawsuits are assigned to various federal units. For more details, consult Tomio (2002b).

[18] The municipalities that underwent a (successful) secession process during 1990 to 1993 were installed in 1993, after the 1992 municipal elections, which defined the mayors for the emancipated localities. Similarly, the municipalities that filed a secession process during 1994 to 1996 were installed in 1997, after the municipal elections of 1996. So, in the 90s, the secessions were only finalized in 1993 and 1997.

Average Area (km²)	1,889.69	1,669.13	1,549.21	1,543.19	1,527.89	-19%

Source: IBGE

The hypothesis formulated by Tomio (2002a) is also helpful in explaining the regional differences that exist in the pace of secession. For example, as the Bahia state governor was contrary to secession and had a solid governing coalition (Tomio, 2002a), no municipalities were created from secession in Bahia in the 1991-2000 period. The quantity of available localities is another factor that explains regional differences in the pace of municipality creation. Another structural characteristic, already discussed by Gomes and MacDowell (2000) and Boueri et al. (2013), is that secession favors the creation of small and micro municipalities. For example, of the 1,016 new municipalities formed from 1991 to 2000, 976 (96 percent) had populations of less than 20,000 inhabitants, and 560 (55.1 percent) had populations of less than 5,000. The sharp decline in average population of municipalities was accompanied by one in geographical area during the period.

Through Table 1, it can be seen that after 1997 there was relative stability in the number of municipalities in the country. This fact is a consequence of Constitutional Amendment No. 15 in 1996, which changed Article 18 of the Federal Constitution to impose stricter rules for district emancipation in order to slow the multiplication of municipalities. In short, the states continue to legislate in matters relating to municipal dismemberment, but now only in periods determined by federal law. In addition, Constitutional Amendment 15 (CA 15/1996) established the need for a plebiscite for all people concerned (defined as the entire population of the original municipality) and required that financial feasibility studies be submitted by the districts interested in secession. However, until the present date, this federal law has not yet been approved by the president[19], and thus the emancipation of districts is legally prohibited[20]. The city establishments that occurred in 1997 were only achievable because the processes were initiated before CA 15/1996. The 58 municipalities created between 2001 and 2010 are the object of legal conflicts, so that even the Supreme Federal Court (STF) considers them illegal (Acayaba, 2008). CA 15/1996 suffered criticism from politicians and academics favorable to the secession of municipalities. One of the main arguments is that the ability to achieve economies of scale in the provision of public goods is

[19] In November 2013 and August 2014, projects to regulate the creation, merging and dismemberment of municipalities were approved by the Senate. However, President Dilma Rousseff fully vetoed both projects (Haubert and Guereiro, 2014).

[20] Thus, from 1997 on, the Brazilian situation resembles that found by Epple and Romer (1989) for US municipalities, where institutional rules tend to make detachments difficult.

associated with the degree of cooperation of a municipality with its neighbors, and not necessarily with the size of its geographical border (G. Bel & Fageda, 2006). Note that, in this perspective, inter-municipal cooperation is clearly viewed as fundamentally motivated by the possibility of obtaining public services and goods at lower costs, i.e. a way for obtaining scale economies (Breuillè et al., 2011).

Given these institutional circumstances, it is pertinent to ask what leads particular districts to request emancipation. The literature suggests some principles. Brink (2004), through a theoretical model of the median voter, analyzed the political and economic conditions that led individuals to choose to separate their city from the larger territory. The author showed that the population of a specific area of the municipality with distinct political preferences, a greater size or a higher level of income in relation to the rest of population is more likely to opt for secession. In addition to these general motivations, others are more specific to the Brazilian case. Bremaeker (1993) argued that the main reason for the emancipation of districts in Brazil is neglect (in terms of public service) by the municipal administration. Gomes and MacDowell (2000) and Boueri et al. (2013) suggested that the rules for the transfer of federal funds are what really encourage municipal secession. It is known that the federal government transfers a portion of federal tax funds to municipalities through the Municipal Participation Fund (FPM), which is the main source of revenue of most Brazilian municipalities. Considering the rules for FPM distribution among the municipalities, there is at least one that generates large inequities: the fixed minimum quotas according to population size. A municipality that has a population well below a specific limit will receive the same amount of funds as one that has a population close to it. For example, a municipality with 10,188 inhabitants (boundary of the first population group) receives the same amount of FPM resources as a municipality with 805 inhabitants. This creates a strong incentive for secession: a small city of 10,000 inhabitants that is divided will (together) receive twice the FPM resources in the post-separation period.

Although the debate about the motivations for municipal secession (or district emancipation) is an important one, this study does not delve into the matter. Our interest is the post-secession period. In particular, we want to investigate the relationship between secession and the fiscal behavior of the municipalities involved.

3.3. Empirical Strategy

Considering that we are interested in measuring the causal impact of secession on municipal public expenditures, the ideal strategy would be to compare what the municipalities that experienced secession spent to what they would have spent if the secessions had not occurred. However, it is impossible to get such counterfactuals. So we use a quasi-experiment approach and consider the Difference-in-Differences estimator (DiD) applied to annual data from 1991 to 2000. This estimator seeks to compare the change in the outcome of the treated group (municipalities that experienced some secession) before and after the intervention with the change in the outcome of the control group (municipalities that did not experience secession), in the same period. In this way, we estimate the following Difference-in-Differences specification with municipality-year observations:

$$Y_{it} = \delta Secession_{it} + \theta p_{it} + \mu_i + \lambda_t + t\omega_i + \varepsilon_{it} \quad (1)$$

where y_{it} is the log of public expenditure of municipality i in year t and $Secession_{it}$ is the treatment variable, which assumes *one* for the sectioned localities in secession and post-secession years and *zero* otherwise. The parameter of interest is δ, which measures the average effect of the treatment on the treated units (ATT). Additionally, p_{it} is a control variable that measures the population size; μ_i is the municipality fixed effect; $\lambda_t$isthe time effect; $t\omega_i$ is an interaction between a municipality dummy and a linear time trend; and ε_{it} is the error term. The municipality fixed-effect term captures time invariant differences in the municipalities, such as geographical properties and fixed institutional aspects. The time-specific effects control for time-varying characteristics common to all municipalities, such as macroeconomic shocks. Finally, the municipality-specific trend controls for time-varying unobservables that evolve at constant rates and for particular spending trends. In equation (1), initially we did not include more socio-demographic controls because they are not collected yearly for the Brazilian municipalities; however, the set of control variables is expanded for the robustness checks[21].

To validate the DiD estimator, it is necessary for the outcome variable trends to be the same in the untreated and treated municipalities in the absence of treatment (Angrist and Pischke, 2008). This is the key identification assumption of DiD and is known as the

[21] To check the robustness of the results regarding the addition of the time-varying controls, we estimated (in subsection 5.2) a two-period DiD containing a number of controls with potential to simultaneously affect the treatment and outcome variables. As these socio-demographic controls are collected every ten years through the Brazilian census, the model is estimated for one pre-treatment year (1991) and one post-treatment year (2000).

common-trend assumption. Although it is not possible to test this assumption – since we cannot observe the treated group in the absence of treatment – we can get some indication about this validity by checking the pre-treatment expenditures trends. If the expenditures trends of the sectioned and non-sectioned municipalities evolved in similar ways before the secession, they would be expected to continue to do so in the post-secession periods in the absence of reforms. We will evaluate the pre-treatment trends in two different and complementary ways: the first is through a graphical inspection and the second is through the estimation of a DiD equation containing anticipatory effects. In this sense we follow Autor (2003) and estimate the parameters of the following dynamic specification:

$$y_{it} = \sum_{\tau=1}^{2} \beta_{\tau} Secess_{i,t+\tau} + \sum_{\tau=0}^{3} \beta_{-\tau} Secess_{i,t-\tau} + \theta p_{it} + \mu_i + \lambda_t + t\omega_i + \varepsilon_{it} \tag{2}$$

This equation is estimated using three kinds of treatment variables. Firstly, there is *anticipatory effects* (also known as *lead dummies*), denoted by $Secess_{i,t+\tau}$. This variable only assumes *one* for the sectioned municipalities in years before the secession. Second, there is *post-treatment effects* (also known as *lag dummies*), denoted by $Secess_{i,t-\tau}$. It only assumes *one* for the treated municipalities in years after the secession. Finally, the *contemporary effects*, denoted by $Secession_{i,t}$, assumes *one* in the secession year for the sectioned localities. Since the secessions occurred in 1993 and 1997, and our database contains annual information from the 1991-2000 period, equation (2) is estimated using three post-treatment effects, two anticipatory effects and one contemporary effect ($\tau = 0$).

It is important to note that if the anticipatory coefficients (β_{τ}) are statistically significant, then there were differences between the public expenditures of the sectioned and non-sectioned municipalities even before the secession process, a situation that weakens the common-trend assumption. The post-treatment dummies of equation (2) allow checking whether the secession impact grows or fades as time passes (Angrist & Pischke, 2008), a question of substantial interest. A plausible hypothesis is that in the year of the secession, the created municipalities have a big rise in public expenditures due to their need for new buildings, installations for the public administration, and the purchase of capital-intensive goods. In subsequent years, with less need for new investments, these public expenditures tend to decrease.

Based on the discussion in section 2, specific concerns should arise about potential for the generation of bias in the process of the secession of municipalities. As was discussed, a secession can fail in two ways. The first is when the concerned population votes against the proposal in the plebiscite, keeping it from even getting to the state legislature. The second is when either the legislators or governor block it during the legislative process. Specifically, if it passes the assembly, it can be vetoed by the governor. If this happens, there are political interactions between the legislature and the state executive that determine the success or failure of the law to create a new city (Tomio, 2002a). These interactions are related to party coalitions, political ideologies, the lobbying power of the postulant district and expected political benefits.

Thus, in order to reduce or eliminate the endogeneity and strengthen our causal claim, in addition to considering municipalities fixed effects, time-specific influences, and a time trend, in our benchmark specification we also consider only municipalities that have tried a boundary reform (treated and untreated units). In other words, we use a control group formed only of municipalities that attempted secession, but were not successful, and call it *almost treated*. The strategy is also supplemented with robustness checks considering other kinds of control groups (based on neighboring municipalities and on previous boundary reforms) and the inclusion of more socio-economic and demographic variables as controls (available from demographic censuses). For comparison purposes, we also present estimates of the parameter of interest in equation (1) using the control group formed of all untreated municipalities, i.e., using the full sample.

Finally, in order to compare municipalities in different periods, the present study will utilize Minimum Comparable Areas (MCAs) as the observational units: these are areas with constant borders over time (Reis, Pimentel, & Alvarenga, 2008). Thus, if there was secession, the MCA of the pretreatment period corresponds to the original municipality, and the MCA of the post-treatment period is the sum of the two revised municipal territories, i.e., the total area under consideration does not change between the pre- and post-treatment times. If there was no secession, the MCA is exactly equal to the municipality in both periods. The strategy also addresses more complex secessions, such as the merging of parts of different municipalities to form a single new municipality[22]. Note that this strategy is analogous to the ones that have studied the impacts of municipal amalgamations on local public spending; but, here, instead of aggregating municipalities' spending in the initial period, we do it in the final (post-treatment) periods. Here, it is important to note that the strong economic dependence of small municipalities on fund transfers from the federal and state governments, and the

[22] On the following pages, we will use the term *municipalities* as a synonym for MCAs.

municipalities' recent administrative autonomy (provided for by the Constitution of 1988) make municipal tax competition certainly less important in Brazil[23].

3.4. Data and Descriptive Statistics

Aiming to analyze the relationship between municipal secession and public expenditures through a Difference-in-Differences model (equations (1) and (2)), we built a set of municipality-year panel data containing information for the 1991-2000 period. This time interval allows us to measure the long-term effects of secession, which is feasible since there is a strong shift in the structures of subnational units (Fox & Gurley, 2006). We used data from the 4,267 municipalities, of which 652 (15.28 percent) underwent a boundary reform, i.e., a secession, in their territories. Of those that did, 266 (40.4 percent) went through their secession in 1993, 314 (48.6 percent) in 1997 and 72 (11.04 percent) in both years. In relation to the control group composed of the almost treated municipalities, the information about the districts that tried an emancipation process unsuccessfully was obtained from the work of Tomio (2002b) and through the records of consultations in the legislative assembly of each Brazilian state[24]. From these sources, we obtain 325 almost-treated municipalities.

Data relative to municipal expenditures were obtained from the Secretary of the National Treasury (STN), which is subordinate to the Finance Ministry. We used two main kinds of expenditures: current expenditures, referring to the costs of maintenance and operation of public services (like payrolls and administrative costs), and capital expenditures, referring to purchases of machinery, vehicles, buildings and the like. On average, 79 percent of municipal expenditures are current expenditures. We could analyze more disaggregated expenditures, such as education and health expenditures, legislative costs and payrolls; however, in the Brazilian case, adopting this approach has shortcomings, since these expenses are tied to certain revenues and/or have constitutional limits, making them almost exogenous to the municipality's control. For example, Article 212 of the Brazilian Federal Constitution

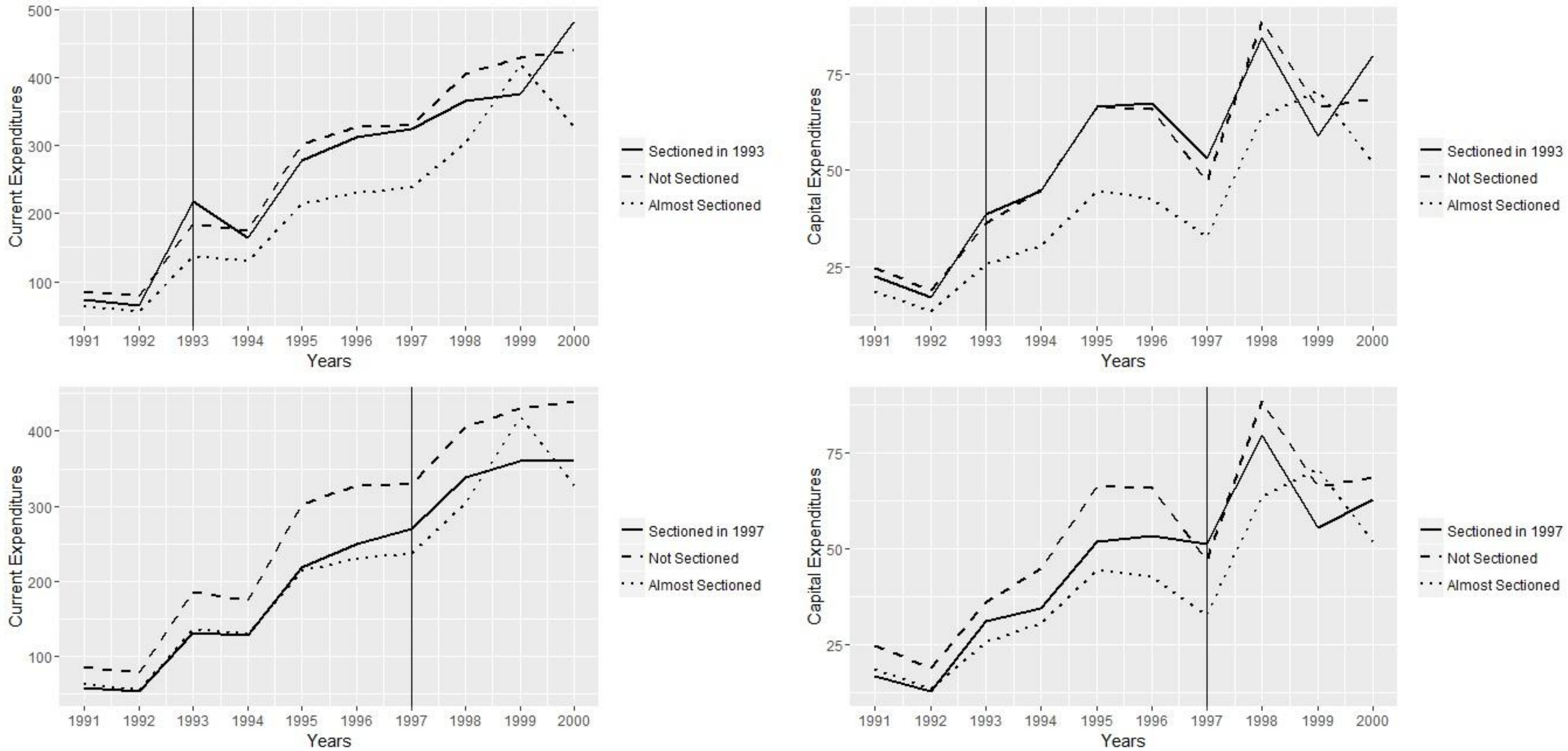
[23] In addition, there is also a juridical argument that makes municipal tax competition less effective in Brazil; as argued, for example, by Sá (2011), the jurisprudence established by the Superior Tribunal of Justice (STJ), the second-highest Brazilian court, is that the municipal ISS taxes have to be paid in the locales where services are rendered, not in the locales of residence of the firms.

[24] We only have information about the districts that tried emancipation for the following states: São Paulo, Rio de Janeiro, Espírito Santo, Minas Gerais, Paraná, Santa Catarina, Rio Grande do Sul, Bahia, Pernambuco and Paraíba. Therefore, in the estimates involving the almost-treated control group, we only use the treated MCAs from these states. Of the 4,267 MCAs of the country, 2,955 (69.25 percent) are located in these states.

states that at least 25 percent of municipal revenues (obtained via own-taxes or transfers) should be relegated to educational activities. In addition, Article 29 provides that expenditures for the payrolls of municipal councilors shall not exceed the limit of 5 percent of municipal revenues.

Figure 1 presents the annual evolution of both types of expenses (in terms of municipality means) for treated (sectioned), untreated (non-sectioned) and almost-treated (attempted sectioning but failed) municipalities. The two graphs on top compare the municipalities seceding in 1993 with those that never did it. The two graphs on the bottom are analogous, but the shock occurs in 1997. It can be seen that the trajectories of the expenditures of the non-sectioned and sectioned groups) are quite similar in all cases, which suggests the common-trend assumption holds.

Figure 1: Evolution of Current and Capital Expenditures by Municipality Status (1991-2000)



Note: The mean values (by municipality) for capital and current expenditures were deflated for R$ of 2000. These expenditure items are in per capita terms, and the population estimates for each year were obtained from IPEADATA. The solid lines represent the average expenditures of those municipalities that were sectioned. The dashed lines indicate the never sectioned municipalities. The dotted line represents the almost treated municipalities.

As discussed in section 2, for robustness purposes, we will do a two-period Difference-in-Differences estimate containing a wide variety of time-varying controls with potential to simultaneously affect the treatment and outcome variables. The first set of these controls corresponds to socioeconomic and demographic variables: per capita income, population, employment rate (ratio of employed population to economically active population), urbanization rate (ratio of population living in urban areas to total population), elderly population (proportion of population over 60 years old), recent immigrants (ratio of five-years-or-fewer immigrants to total population) and the Gini coefficient to measure income inequality. These variables were constructed using the 1991 and 2000 Brazilian Demographic Censuses obtained by IBGE. The second set of controls includes institutional and political characteristics of the municipalities that were compiled by the Superior Electoral Court (TSE). Thus, the following variables were calculated for each MCA: (i) political participation of residents in the elections, measured as the ratio of voters who attended the first round of general elections to the total electorate; and (ii) political divergence among the population, measured as the proportion of valid votes not cast for the most-voted-for candidate for president in the first round of general elections (relative to that municipality). Unfortunately, as there is no digital data for municipal elections prior to 1994, we utilized data from the general elections of 1989 (pretreatment period) and 1998 (post-treatment period). In addition to these variables, we used the revenues of the Municipal Participation Fund (FPM)[25] as an additional control to investigate the mechanisms behind the relationship between secession and municipal expenditures. The variables that are in monetary units were deflated by the INPC (National Consumer Price Index) for R$ of 2000.

Table 2 presents descriptive statistics (mean and standard deviation) for the total sample of municipalities and for the treated, untreated and almost-treated subsamples in the pre-intervention year of 1991. Additionally, mean difference tests between the treatment group and controls groups are reported. Some numbers of Table 2 should be highlighted. First, there are significant differences between the treated and not-treated municipalities, a natural consequence of the non-randomness of the treatment. In general, the municipalities that underwent secession are characterized by having less government participation, measured by lower levels of public expenditures and FPM revenue.

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[25] On average, about 48 percent of municipal revenues come from the FPM, while only 5.7 percent are own-revenues.

Table 2 - Summary Statistics – 1991

	All	Treated	Not Treated	Almost Treated	Diff. 1	Diff. 2
Capital Exp. (R$)	23.818	19.135	24.673	18.730	-5.53***	0.405
	(25.77)	(18.66)	(26.78)	(27.98)		
Current Exp. (R$)	82.730	64.192	86.114	63.161	-21.92***	1.031
	(54.30)	(35.71)	(56.39)	(44.76)		
Income pc (per 1,000)	5.899	6.944	5.710	6.207	1.23	0.736
	(68.73)	(18.07)	(74.28)	(29.09)		
Population (per 1,000)	34.409	52.367	31.17	46.55	21.19***	5.814
	(193.19)	(92.80)	(20.60)	(96.50)		
Income Inequality	0.530	0.536	0.529	0.541	0.007***	-0.005
	(0.07)	(0.05)	(0.07)	(0.06)		
Employment Rate	0.959	0.960	0.959	0.953	0.000	0.007***
	(0.04)	(0.03)	(0.04)	(0.03)		
Urbanization Rate	0.536	0.521	0.539	0.510	-0.018*	0.011
	(0.23)	(0.23)	(0.23)	(0.25)		
FPM Revenues (R$)	44.893	29.684	47.685	29.725	-18***	-0.041
	(33.78)	(13.83)	(35.57)	(21.35)		
Voter Turnout	0.834	0.820	0.837	0.835	-0.01***	-0.01***
	(0.10)	(0.12)	(0.10)	(0.09)		
Political Divergence	0.533	0.525	0.534	0.527	-0.009*	-0.002
	(0.12)	(0.12)	(0.12)	(0.11)		
Elderly Population	0.076	0.071	0.077	0.075	-0.006**	-0.004*
	(0.02)	(0.02)	(0.02)	(0.02)		
Recent Immigrants	0.114	0.114	0.114	0.088	0.000	0.026**
	(0.090)	(0.090)	(0.08)	(0.06)		
Number of Municipalities (MCAs)	4267	652	3615	325		

Note: Standard deviations are in parentheses. Capital expenditures, current expenditures, income and FPM revenues are in per capita terms. Diff1 refers to the mean differences between treated and untreated. Diff2 refers to the mean differences between treated and almost treated. *** p<0.01, ** p <0.05, * p <0.1.

This fact is in line with the Bremaeker (1993) hypothesis, according to which the main motivation for municipal secession is the absence of participation of the local administration in the provision of public goods and services. However, considering the almost-treated municipalities all but erases the fiscal disparities between the municipalities. Indeed, the matching between the treated and untreated municipalities becomes much better when we consider this control group. In fact, some divergences are maintained (sectioned localities had a higher employment rate, lower voter turnout, lower elderly population and a higher number of recent immigrants), but they are not generally substantial.

3.5. Results

As argued in the initial section, it is likely that the process of municipal secession in Brazil, by reducing the size of the municipalities, causes reductions in economies of scale for public goods and services, and thus raises public expenditures per capita. In this section, we provide a set of evidence in order to analyze this possibility. In subsection 3.5.1 we investigate if the secession of Brazilian municipalities indeed generates increased per capita expenditures, considering both current and capital expenditures. In subsection 3.5.2 we present a set of robustness tests to check the sensitivity of the results. Finally, in subsection 3.5.3 we analyze different potential mechanisms linking secession to variation in municipal expenditures and discuss if our results can actually be associated with reduced economies of scale.

3.5.1. Does secession increase per capita expenses?

Table 3 shows the benchmark results for equation (1). The outcome variables are the logarithms of the current and capital expenditures per capita. Panel A refers to the first category and Panel B to the second. As discussed in section 3, we will use two control groups: those formed of all municipalities who have not undergone secession (full sample, columns (1) and (2)) and another more restricted one formed of only the municipalities who tried and failed to accomplish boundary reform (almost treated, columns (3) and (4)). Moreover, equation (1) is estimated with and without the municipality-specific trend.

Table 3: Effects of Secession of Municipalities on Public Expenditures: Basic Specifications.

Panel A: Current Exp.	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
Secession	0.101***	0.053***	0.057***	0.078***
	(6.73)	(2.71)	(2.97)	(3.46)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Municipality-Specific Trend	No	Yes	No	Yes
N. Observations	39,561	39,359	6,982	6,957
Panel B: Capital Exp.	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
Secession	0.223***	0.176***	0.185***	0.147***
	(7.89)	(4.22)	(4.52)	(2.64)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Municipality-Specific Trend	No	Yes	No	Yes

N. Observations	39,526	39,325	6,979	6,954

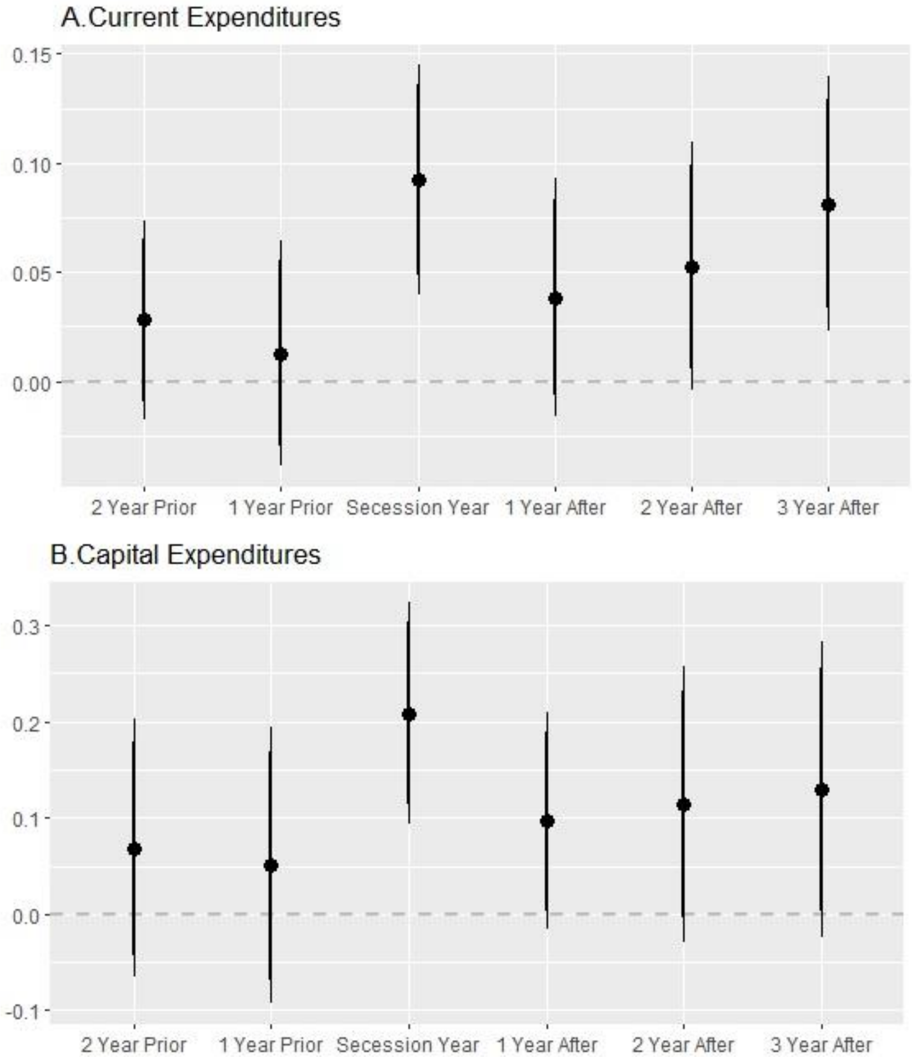
Note: *** P <0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. T-values are in parentheses. The log of the population size is included as a control in all specifications.

Initially, we note that the variable that measures the impact of secession on public expenditures (the ATT) reveals that the municipalities that experienced secession increased their current expenditures per capita by about 10.1 percent compared to the situation of not experiencing secession. This effect is substantially higher (22.3 percent) when considering capital expenditures. This evidence contradicts the arguments that suggest that the decrease in the size of the subnational units leads to lower levels of public expenses, either by closer ties between the politicians and voters – which favor efficiency – or by increased competition between municipalities (Nelson, 1992; Swianiewicz, 2010).

A common concern in the basic specificationis that the municipalities with rising expenditures are the most likely to secede. If this is the case, a positive impact for secession might simply be spurious, i.e., just reflect a difference in expenditures trends. To evaluate this issue, we add the municipality-specific trend in equation (1), (column (2)). Although there is some reduction in the value of the estimated coefficient, we note that the incluson of this control does not qualitatively change our estimates, revealing that the secession process is not simply determined by the trends of public spending. Additionally, when we restrict the control group to the almost-treated ones (columns (3) and (4)), the secession impact remains positive and statistically significant, albeit with a decrease in the coefficient magnitudes. These evidences point to ATT estimates that are also robust for unobserved variables affecting the aspiration for secession, making our results more credible. Therefore, in our most reliable estimates (column (4)), the ATT indicates that the sectioned municipalities increased their per capita current expenditures by 7.8 percent and capital expenditures by 14.7 percent. Note that the former value is similar, and the latter one higher but not so distant from the relative reductions of expenditures experienced by some developed countries after an amalgamation process (Reingewertz, 2012; Blesse & Baskaran, 2016).

As discussed in section 3, for the common-trend assumption to be valid, it is necessary for the pretreatment trends of the treated and untreated municipalities to be equal. The common trend is suggested by the set of evidence provided in Figure 1. Another way to evaluate this is through the estimation of equation (2). The leads and lags specification is also useful for understanding the time dynamics of the secession process. Figure 2 shows the results of the coefficient estimates from equation (2) considering only the specification that simultaneously includes the municipality-specific trend and the almost-treated control group (the associated table is available in the Appendix).

Figure 2 - Impact of Secessions on public expenditures for years prior, during and after the process, 1991-2000.



Note: The outcome variables are the logarithm of expenditures per capita and the control group is formed by almost treated municipalities. The models include the municipality-fixed effect, time fixed-effect, municipality specific-trend, and log of population size. The dots are the regression coefficients and the vertical lines correspond to 99% (inner line) and 90% (outer line) confidence bands.

Firstly, it is observed that the anticipatory effects are not statistically significant in either specification. Therefore, during the pre-intervention period there were no differences in the expenditure behavior between those municipalities that would and would not experience secession. It should be noted that this result also represents evidence against the existence of some confounding factor that could explain the increased expenditures. If any such factor existed, it would not come into play in the periods of pretreatment, only at the same time as the secession shock, which is unlikely.

In relation to the time dynamics, it is observed that in secession years the affected municipalities increased their per capita current expenditures by 9.3 percent and capital expenditures by approximately 21 percent. These effects fade in subsequent years, especially in relation to capital expenditures, where the impact falls almost 50 percent in the first year after the boundary reform. This evidence is consistent with our initial hypothesis that, due to installation and entrance costs, the capital expenditures tend to be higher in the year that the event occurs and less in subsequent years. Similarly, the nature of current expenditures, which are primarily payrolls with fixed contracts, can explain its greater stability.

3.5.2. Robustness checks

In this subsection, we present a set of robustness checks in an attempt to verify the sensibility of our results: we use different control groups (drop nearby municipalities and the previously sectioned ones) and include a wider set of time-varying covariates obtained from the demographic census data.

Firstly, the sectioned municipalities' neighbors may react to their increase in public expenditures by various types of strategic actions. There may be a boost in public expenses due to the need for complementary public goods that are shared or even as a result of mimicking behavior (Solé-Ollé, 2006; Costa et al., 2015). However, when there is a substitute public goods, the optimal strategy of a nearby municipality is to reduce their public expenditures in a kind of free-rider behavior. Such spatial dependence related to public expenditures tends to cause a bias in our estimates. If the dependence is positive – the nearby city increases spending in response to the first's secession – the ATT will be underestimated. If the dependence is negative, the ATT will be overestimated. To verify if the results obtained above are influenced by this situation, we will drop the contiguous neighbors of the sectioned municipalities from the control group, which leads to the exclusion of 1,859 municipalities. This strategy allows the reduction of bias arising from potential spatial interactions.

Another way to check for bias in the results is by disregarding the secessions that occurred in the 1980s[26]. Specifically, this is done through the use of a potentially cleaner

[26] Between 1980 and 1990, 500 municipalities were created in Brazil. Nevertheless, we do not evaluate the secessions of this period because the institutional context of the country was very different from the present. The

control group formed by removing from the sample all of the municipalities that changed their territories by secession in the period 1980-1990 (ones that could thus present less or more potential for another boundary reform).In our database, there are 625 municipalities in this situation. Removing this specific group maybe important because, if secession implies less efficient spending, they represent a group with relatively high per capita expenditures in the pretreatment period, which, without adequate control, could impose a positive bias on the estimate. Thus, this new control group consists of those municipalities that had no secessions from 1980 to 2000.

Table 4 presents the results for each of these robustness checks. Column (1) shows the estimation results of the coefficients of equation (1) when removing the neighbors of the sectioned municipalities from the control group; column (2) presents the estimate obtained by dropping from the control group the municipalities that experienced secession in the 1980s. The results presented in columns (3) and (4) are the corresponding estimations when using the almost-treated municipalities as the control group.

Table 4 - Effects of Secession of Municipalities on Public Expenditures: Robustness Checks Using Different Control Groups

Panel A: Current Expenditures	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
	Drop Neighbors	Drop 80s	Drop Neighbors	Drop 80s
Secession	0.047**	0.051**	0.076***	0.074***
	(2.39)	(2.63)	(3.23)	(3.24)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Municipality-Specific Trend	Yes	Yes	Yes	Yes
N. Observations	22,439	33,884	5,446	6,364
Panel B: Capital Expenditures	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
	Drop Neighbors	Drop 80s	Drop Neighbors	Drop 80s
Secession	0.147***	0.177***	0.119**	0.143**
	(3.46)	(4.23)	(2.06)	(2.55)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Municipality-Specific Trend	Yes	Yes	Yes	Yes
N. Observations	22,423	33,857	5,443	6,364

Note: *** p<0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. The t-values are in parentheses. The outcome variables are the logarithms of the current and capital expenditures per capita. The log of the population size is included as a control in all specifications.

The set of evidence presented in Table 4 shows that our results are robust to the potential presence of spatial interactions of public expenditures and to the presence of

rules of boundary reforms were defined by a constitution created during a military dictatorship (1967), and there was greater administrative centralization with little autonomy for the subnational governments.

previously sectioned localities. The results are very similar to those obtained in Table 3. Considering the results of the current expenditures (Panel A), it can be seen that the estimated ATT is practically unchanged. In relation to the capital expenditures (Panel B), it is observed that the estimated ATT falls marginally when the neighbors of the sectioned municipalities are dropped.

In addition to investigating the sensitivity of the results of different subsamples (Table 4) and to the existence of a pre-trend of increased expenses (Figure 2), we also tested the robustness of our results with the inclusion of more time-varying confounders. The omission of variables that simultaneously affect the secession and public expenditures could potentially generate endogeneity in our estimates, inhibiting a causal interpretation. For example, it may be that the richer or more unequal municipalities are precisely those who have sectioned, and thus our estimated ATT does not capture the secession effect, but just a socio-economic difference between localities. Although the use of a control group formed only by the almost-treated municipalities reduces this concern, an alternative approach is to include socio-economic controls in equation (1). However, as the municipal socioeconomic variables are collected only every ten years, we must restrict specification (1) to a two-period DiD, containing one pretreatment period (1991) and one post-treatment period (2000). The new set of control variables contains the municipalities characteristics previously presented in Table 2: per capita income, employment rate, urbanization rate, voter turnout, population, political divergence, percentage of elderly population and percentage of recent immigrants. The results of these new estimations are reported in Table 5[27].

Table 5 - Effects of Secession of Municipalities on Public Expenditures: Robustness Checks Including Socio-Economic Controls

Panel A: Current Expenditures	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
Secession	0.177***	0.166***	0.044	0.050*
	(5.93)	(5.58)	(1.47)	(1.70)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
N. Observations	8,211	8,193	1,436	1,431
Panel B: Capital Expenditures	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
Secession	0.318***	0.313***	0.248***	0.250***
	(5.57)	(5.56)	(4.13)	(4.17)

[27] Full results are available upon request.

Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
N. Observations	8,200	8,182	1,434	1,429

Note: *** p<0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. The t-values are in parentheses. The outcome variables are the logarithms of the current and capital expenditures per capita. The results were obtained through a two-period Difference-in-Differences estimation.

The numbers in Table 5 indicate that our results are robust to the inclusion of more socioeconomic variables. Comparing column (3) with column (4), there is only a marginal change in the estimated ATT with the inclusion of the time-varying controls. Additionally, although the estimated coefficients of a two-period DiD using the full sample of control municipalities (columns (1) and (2)) are much higher than those obtained previously, the coefficients obtained using the almost-treated control group (columns (3) and (4)) are more similar to the corresponding values reported in column (3) of Table 3. This reveals that the estimation using the almost-treated control group is more robust to model miss-specification.

3.5.3. Understanding the mechanisms relating secessions to per capita expenditures.

In this subsection, we analyze the possibility that the positive relationship that we found between municipal secession and higher levels of expenses can be explained by others channels than the loss of economies of scale. In addition, we analyze the information about inter-municipality cooperation to detect potential heterogeneities in public-expenditure behavior according to the condition of belonging or not to an inter-municipal consortium.

More specifically, we first consider three other mechanisms that can potentially explain the positive relationship between municipal secession and higher levels of expenses: (i) increased revenues, since one of the main motivations for secession was the potential increase of federal transfers via FPM (Gomes & Mac Dowell, 2000); (ii) improvements in the provision of public goods and services (Reingewertz, 2012; Allers & Geertsema, 2016); and (iii) corruption associated with the secession process.

A possible way to test the first pathway is to include revenues from FPM in equation (1) as an additional control. If the coefficient that measures the ATT remains statistically significant and positive, the evidence is that, regardless of the rise in revenues, the municipalities that had experienced secession increased their expenditures. It is worth noting that the FPM revenue is a potential outcome of the treatment variable, so to add it to equation (1) generates bias in the coefficients, i.e., revenues are considered bad controls (Angrist &

Pischke, 2008). The inclusion of bad controls in a regression model is only useful for investigating the mechanisms that explain a relationship between two variables (see, for example, Maccini & Yang, 2009). Table 6 reports the results when the log of the FPM revenues is included as an additional control in equation (1).

Table 6 - Effects of Secession of Municipalities on Public Expenditures: Specifications with FPM Revenues per Capita Added

Panel A: Current Expenditures	(1)	(2)
	Full Sample	Almost Treated
Secession	0.012	0.025
	(0.88)	(1.63)
Log (FPMpc)	0.519***	0.636***
	(8.53)	(8.33)
Municipality FE	Yes	Yes
Time FE	Yes	Yes
Municipality-Specific Trend	Yes	Yes
N. Observations	39,359	6,957
Panel B: Capital Expenditures	(1)	(2)
	Full Sample	Almost Treated
Secession	0.130***	0.089*
	(3.42)	(1.74)
Log (FPMpc)	0.594***	0.720***
	(9.76)	(8.44)
Municipality FE	Yes	Yes
Time FE	Yes	Yes
Municipality-Specific Trend	Yes	Yes
N. Observations	3,925	6,954

Note: *** p<0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. The t-values are in parentheses. The outcome variables are the logarithms of the current and capital expenditures per capita. The log of the population size is included as a control in all specifications.

The numbers in Table 6 indicate that, after controlling for per capita FPM revenues, the impact of secession on the current expenditures become statistically insignificant. This result suggests that local current expenditures only increase due to higher federal transfers. Nevertheless, the impact of secession on the capital expenditures, despite some reduction, remains positive and statistically significant. This indicates that the mechanisms that explain the raised levels of capital expenses in the wake of secession are not simply related to the increase of FPM resources.

We now consider the possibility that increased expenditures are associated with better provision of public services and goods by those municipalities that experienced secession. There are some problems in testing this channel. First, it is difficult to establish a causal link between secession and the provision of public goods and services. Second, inasmuch as the public goods and services are supplied by three levels of government in cooperation (federal, state and municipal), it is difficult to isolate those exclusive to the municipalities. Third, as argued by Reingewertz (2012), not all municipal services can be quantitatively measured.

However, despite these challenges, we test the effect of these secessions on the supply of some public services in the hope of getting some additional information. A positive impact would be an indication that the increase in capital expenditures should not necessarily be associated with a loss of economies of scale, but with the better provision of public services. The approach is thus similar to that of Mattos and Ponczek (2013). Specifically, we tested for the impact of secession on the provision of the following public services: garbage collection, access to electricity, running water, public sewer network (all measured as the percentage of houses with the services in question), and school attendance of the children between 7-14 years old[28]. Only garbage collection can be considered as the sole responsibility of the municipalities, but since municipalities have exclusive jurisdiction over their kindergartens and elementary schools, the school attendance of children 7-14 years old (elementary-school level), might be a good way to measure the provision of education by the municipalities. Furthermore, expansion of the other services by the state governments generally involves additional local expenditures on urban infrastructure by the municipalities. Because these variables are only collected every ten years, we estimate two-period DiDs, as in Table 5. The results are summarized in Panel A of Table 7.

It should be emphasized that, because these variables do not cover the entire output of the municipalities, there is of course the risk of the enhancement of other public services not considered. An indirect way to assess whether there was public goods improvement is to check if there were changes in the attractiveness of the sectioned cities. Improving the public service levels or quality would make a municipality more attractive to live in, which we would expect to be capitalized in house prices or house stock (Oates, 1969; Reingewertz, 2012; Allers & Geertsema, 2016). Thus, we also estimate the impact of secessions on the local housing market. Using data from the Institute of Applied Economics Research (IPEA), Panel B of Table 7 shows the results for two outcomes: the log of the number of houses and the log of the value of the residential capital stock[29].

As can be seen in Panel A of Table 7, our new set of evidence does not indicate any improvements in the supply of public services associated with the process of municipal

[28] These variables were calculated for the years 1991 and 2000 and are available in the Brazilian Demographic Census.

[29] Property prices are not collected for all municipalities in the time period considered in our analysis.

secession in Brazil. In fact, according to Mattos and Ponczek (2013), it turns out that secession is associated with a deterioration in the supply of some public services. However, apart from the negative association with school attendance, the impacts are not statistically significant, indicating that they are not economically substantive. Hoyt (1991) showed that an increase in the number of jurisdictions generates lower levels of tax rates and public services, due to the potential increase in tax competition among the local units. According to this perspective, the absence of effects on public services in the Brazilian case (Table 7) is consistent with the absent or weak fiscal competition among Brazilian municipalities, as discussed in section 2. This is reinforced by the estimate of no impact by the secessions on local tax revenues (see Table A2 of the Appendix).

Table 7 - Effects of Secession of Municipalities on Supply of Public Services and on Housing Market

Panel A: Public Services	(1)	(2)	(3)	(4)	(5)
	Garbage	Electricity	Running Water	Sewage	School Attend.
Secession	-0.018	-0.014	-0.008	-0.022	-0.054***
	(-1.01)	(-1.26)	(-0.85)	(-1.52)	(-5.07)
Municipality FE	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes
N. Observations	1,457	1,457	1,457	1,457	1,457
Panel B: Attractiveness	(1)	(2)			
	Houses	Residential Capital			
Secession	0.010	-0.038*			
	(0.91)	(-1.93)			
Municipality FE	Yes	Yes			
Time FE	Yes	Yes			
Controls	Yes	Yes			
N. Observations	1,457	1,457			

Note: *** p<0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. The t-values are in parentheses. The control group is formed of the almost-treated municipalities. The results were obtained through a two-period Difference-in-Differences estimation.

A similar result was found for the variables that measure the dynamics of the housing market (Panel B). It is observed that there was not an increase in housing stock or an increase in the value of residential capital in the sectioned municipalities. Actually, taken together, these results suggest that there was a decline in the property prices of the localities that underwent secession. Therefore, at least for the two variables we used to reflect municipalities' attractiveness, the evidence of Table 7 does not support a possible improvement in the provision of public services in the municipalities that participated in a secession process.

The third potential channel we examine that might link higher capital expenditures to the secession of municipalities is corruption and rent-extracting behavior. Due to the huge inflow of federal grants in the sectioned municipalities, politicians and local leaders could be encouraged to engage in corrupt and inefficient actions. For example, the additional resources may be diverted for personal benefit or contracts might get padded with exorbitant fees. Under such circumstances, the higher expenditures are the consequence of local rent seeking, not of a reduction in economies of scales or improvement of public services. Investigating this mechanism is empirically difficult, since measurements of rent seeking are not directly available and the recognized corruption variables in Brazil do not cover our sample of municipalities. Furthermore, it is clearly possible that this hypothesized reason for capital spending increases would act together with loss of scale economies. So here we rely on previously collected indirect evidence about the election process that is useful for generating insights concerning the mechanism's validity.

In this respect, as convincingly shown by Ferraz and Finan (2011), the risk of being directly exposed and judged by the population through the recent institution of incumbent re-elections in municipalities have reduced the levels of corruption. Noting that incumbent re-election was instituted in 2000, we identified that, while 65.9 percent of the mayors of municipalities without any kind of boundary reform ran for re-election in 2000, this percentage was 73.2 percent for those in the municipalities that undertook boundary reforms. Thus, the motivation against corruption provided by the desire for re-election was more present in the set of municipalities that had experienced some kind of secession than in those municipalities without any kind of boundary reform. On the other hand, supporting our concerns, Brollo et al. (2013) recently provided evidence that higher intergovernmental grants to Brazilian municipalities result in an increase in corruption. Thus, based on the available evidence, we conclude that it is probable that rent-seeking gains as well as losing scale economies have affected capital spending in Brazilian municipalities experiencing secession.

Besides investigating the above three mechanisms, we also evaluate if the trajectory of capital expenditures behaves differently among Brazilian municipalities with or without some form of inter-municipal cooperation. It is broadly recognized that one of the prior objectives of municipal cooperation was to provide public services and goods at lower cost (Agranoff & McGuire, 2004; Bel & Fageda, 2006; Hulst & Van Montfort, 2007; Breuillè et al., 2011; Bel et al., 2013), i.e. to obtain economies of scale. In cases of highly fragmented territories such as that of France, for example, with more than 36,000 municipalities, Breuillè et al. (2011)

highlighted that 95.5 percent of the municipalities were engaged in some form of inter-municipal cooperation. Our new exercise can potentially thus bring some additional light about the economic consequences of secession in Brazil: if losses of scale economies are behind the increases of per capita capital expenditures we have identified for Brazilian municipalities engaged in secession, and if inter-municipal cooperation is mainly motived by the possibility of scale-economy gains, we would expect the municipalities involved in some form of inter-municipal cooperation to exhibit less (or no) rise in spending than those not so involved. On the other hand, the absence of heterogeneity in the impact of municipality secession on the dynamic of capital spending between the two types of municipalities provides some support against the argument.

In Brazil, there exists some form of inter-municipal cooperation data before the 1988 constitution, but much more was collected after that time (Teixeira, Bugarin, & Dourado, 2006; Linhares, 2011). These inter-municipal arrangements, known as inter-municipal consortiums, are nowadays regulated by a federal law (11,107/2005), and generally constitute horizontal, voluntary contracts, with the aim of supplying some type of public service or good (Spink, 2005; Linhares, 2011). According to the Municipality Information Research (MUNIC, 1999), about 41 percent of Brazilian municipalities were part of some inter-municipal consortium in the following areas: health (37.3 percent), education (4.1 percent), housing (0.85 percent), machinery and equipment (4.3 percent), water supply (2.8 percent), sewer (1.5 percent) and garbage collection (3.28 percent). The available evidence and analyzes reveal some consensus about the motivation behind the Brazilian inter-municipal cooperation: given the large numbers of poor and small municipalities, on one hand, and the new municipal competencies specified by the 1988 constitution, on the other, inter-municipal cooperation is a way to guarantee the provision of public services and goods at affordable costs (Spink, 2005; Teixeira, Bugarin, & Dourado, 2006; Linhares, 2011)[30].

Thus, in order to verify if the impacts of secession are heterogeneous among municipalities with and without some form of inter-municipal cooperation, we estimate equation (1) for two subgroups of municipalities: those who participated in some sort of consortium and those who did not[31]. The results are shown in Table 8.

[30]As argued by Teixeira et al., 2006, for example, in some sectors such as health care, large investments are necessary not only in physical capital but also in human skills. Maybe this can explain why the majority of municipal consortiums are about health services.

[31]The information about the municipalities participating in consortiums was obtained through the Survey of Basic Municipal Information (MUNIC) of 1999.

The numbers in columns (3) and (4) of Table 8 indicate that the impact of secession is about thirty percent lower in those municipalities that are involved in a cooperation scheme than in those without any such involvement. This also suggests that the behavior of capital expenditures depends on the potential for economies of scale in the cities. Thus, inter-municipal consortiums appear to be an effective strategy to mitigate the drawbacks from the secession reform.

Table 8 - Heterogeneous Responses by Inter-Municipal Consortiums

Capital Expenditures	(1)	(2)	(3)	(4)
	Full Sample		Almost Treated	
	Consortium	No Consortium	Consortium	No Consortium
Secession	0.148***	0.208***	0.119*	0.154*
	(2.90)	(2.92)	(1.78)	(1.72)
Municipality FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Municipality-Specific Trend	Yes	Yes	Yes	Yes
N. Observations	17,858	21,668	4,050	2,929

Note: *** P <0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. T-values are in parentheses. The log of the population size is included as a control in all specifications.

Consistent with this argument, our data indicate that while 59.2 percent of the municipalities that experienced some kind of secession decided to create some type of inter-municipal cooperation, only 41.1 percent of the municipalities that did not experience any kind of secession did so. More specifically, Teixeira et al., (2006) argued that the ability to provide specialized health services (the sector presenting the highest share of total inter-municipal cooperation agreements in Brazil) commonly requires scales of production that are incompatible with the small local populations in most Brazilian municipalities. Thus, we believe that an important motivation behind the Brazilian inter-municipal cooperation agreements is the requirement for economic efficiency through economies of scale.

It is important to recognize, nevertheless, that it is still possible that cooperation makes it more difficult for local administrators to collect rents, as there are more parties involved in the public expenditures. This possibility makes this evidence also consistent with the existence of rent-seeking gains associated with the secession of Brazilian municipalities.

In sum, the results found in this subsection indicate that the rise in capital expenditures attributable to the secession of municipalities seems related neither directly to the higher revenues obtained through federal grants (FPM) nor to improvements in the supply of public services. Instead, the increase in the capital spending of the sectioned municipalities appears somehow related to their diminished potential to achieve economies of scale and to rent-seeking behavior by local leaders.

3.6. Discussion and Final Remarks

Although there is a vast body of empirical literature investigating the fiscal impacts of the amalgamations of subnational governments, little is known about the reverse movement: the secession of municipalities. Considering that some recent works have found that larger cities can obtain economies of scale and consequently reduce their level of per capita expenditures (Reingewertz, 2012; Blesse & Baskaran, 2016; Hanes, 2015), it is expected that secessions cause exactly the opposite result; i.e., reducing the size of local governments causes a loss of economies of scale and thus higher per capita expenses. This article evaluates the case of Brazil, where over a period of just ten years (1991-2001) there was a 25 percent increase in the number of municipalities. Thus, besides exploring the case of a developing country, our analysis brings a new point to the debate about the impacts of boundary reforms.

The results indicate that the Brazilian municipality-dividing process has more impact on capital expenditures than on current expenditures. On average, municipalities that were sectioned increased their per capita capital expenditures by 14.7 percent. This result was quite expected since the public services associated with capital goods benefit from larger sizes (Ladd, 1994; Fox & Gurley, 2006), and those that are labor-intensive (more associated with current expenditures) generally present lower fixed costs. On the other hand, we also provide evidence that these higher per capita capital expenditures associated with municipal divisions can hardly be explained solely by higher revenues obtained through federal grants (FPM) or by improvements in the supply of public services.

Through an analysis of impact heterogeneity, we bring additional insights to the boundary reforms. First, the time dynamics of the secessions reveal that the increased spending is more concentrated in the year of the event, and there is a decrease in the growth of capital expenditures in the following years, which is consistent with the need for initial purchases by newly created cities. Second, we show that the secession impact is weaker for members of inter-municipal consortiums.

Our set of evidence suggests that reduced scale economies and rent-seeking behavior are the sources of the capital spending increases. Given the increases in capital expenditures that result from the secessions in Brazil, an apparent mitigation strategy would be to encourage more cooperation among the municipalities in the provision of capital-intensive public services, a strategy that could reduce both the posited loss of economies of scale and the increase of spending associated with rent-seeking behavior (once decisions would involve more parties). For example, neighboring municipalities could share their health services, garbage-collection systems, water supplies and public-transportation systems. However, it is necessary to recognize that the available evidence about the effects of municipal consortiums is mixed. While Charlot et al. (2014) showed that cooperation appears to be an appropriate strategy for reducing tax competition and local tax rates, Frère et al. (2014) provided evidence that inter-municipal cooperation does not necessarily have an impact on the level of public spending. More theoretical and empirical research appear to be necessary for understanding the general implications of municipal boundary reforms and inter-municipal cooperation for local public expenditures.

3.7. References

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3.8. Appendix A- Third Chapter

Table 1A – Impact of Secessions on Public Expenditures for Years Prior to, During and After the Process, 1991-2000

	(1)	(2)
	Current Expenditures	Capital Expenditures
Secession (2 Years Prior)	0.028	0.069
	(1.24)	(1.02)
Secession (1 Year Prior)	0.013	0.050
	(0.50)	(0.70)
Secession	0.093***	0.209***
	(3.51)	(3.61)
Secession (1 Year After)	0.081***	0.097*
	(2.81)	(1.72)
Secession (2 Years After)	0.053*	0.114
	(1.85)	(1.59)
Secession (3 Years After)	0.039	0.129*
	(1.41)	(1.68)
Municipality FE	Yes	Yes
Time FE	Yes	Yes
Municipality-Specific Trend	Yes	Yes
N. Observations	6,982	6,979

Note: ***p<0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. The t-values are in parentheses. The outcome variables are the logarithms of expenditures per capita and the control group is formed of the almost-treated municipalities. The log of the population size is included as a control in all specifications

Table 2A – Effects of Secession of Municipalities on Local Taxes Revenues

Local Taxes	(1)	(2)
	Full Sample	Almost Treated
Secession	0.006	0.093
	(0.11)	(1.41)
Municipality FE	Yes	Yes
Time FE	Yes	Yes
Municipality Especific Trend	Yes	Yes
N. Observations	39,204	6,963

Note: *** P <0.01, ** p <0.05, * p <0.1. We used robust standard errors that were grouped at the municipal level. T-values are in parentheses. The log of the population size is included as a control in all specifications.