Local Retail Prices, Product Varieties and Neighborhood Change

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Abstract

We study how local grocery prices within a city are affected by changes in housing markets. Our empirical strategy exploits an exogenous shift in the spatial distribution of construction activity induced by a large-scale, place-based tax exemption in the city of Montevideo. The estimated elasticity of grocery prices to newly-built residential space lies between -3 and -4%. Using a multi-product model of imperfect competition, we show that this negative effect can result from either an expansion in varieties or firm entry. We report evidence supporting the varieties channel, as well as evidence of changes in the composition of local stores.

Keywords: Retail Prices, Housing Stock, Neighborhood Change.
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1. Introduction

The availability of grocery stores and supermarkets is not homogeneous within cities. Differences in the local consumer base across locations can shape the availability of these stores as well as the prices and varieties of goods sold in them. Therefore, physical changes in neighborhoods that influence this consumer base can affect local retail options. The direction of these changes – as well as their welfare and distributional consequences – will in general depend on the local supply response, both in terms of changes in varieties of goods sold and in the entry of new grocery stores.

In this paper, we study how neighborhood change affects local retail opportunities within cities. Specifically, we test whether large scale development of new housing stock within a city influences the price and varieties of groceries available locally, as well as the density of stores in affected neighborhoods. This is motivated by the notion that residential development can affect incumbent households through indirect channels – i.e. beyond the direct effects on the market for housing services – that are relevant to the debate around the welfare impacts of neighborhood change. The focus on groceries in particular is motivated by the fact that these goods have a large consumer base and represent a larger share of spending by households with relatively lower incomes.

New development can affect the market for groceries because it increases local demand for these goods. In the first place, changes in stock may increase residential density - i.e. the volume of consumers at each location – thus scaling up demand. In the second, new stock can affect neighborhood composition. Previous studies have shown that the age of the housing stock can partly explain the dynamic of neighborhoods' economic status (Rosenthal, 2008; Brueckner and Rosenthal, 2009; Rosenthal, 2020). Newly built units often attracts affluent residents with a high willingness to pay for this type of housing (Brueckner, 2011). Through both channels, the local demand for goods and the demand for different varieties may increase with new housing development.

Estimating how residential development affects local conditions in the market for groceries requires dealing with a reverse causality problem: residential development is shaped by local demand for housing and is therefore influenced by local retail options. In addition, neighborhood characteristics such as accessibility to jobs or local crime rates can affect housing demand and grocery supply conditions. To overcome these problems, we exploit quasi-experimental variation from a major housing policy intervention that induced a large re-location in the development of new stock within the city of Montevideo, Uruguay. The policy provides tax benefits to developers building housing in a pre-defined middle-income area of the city, effectively subsidizing development in those locations. Developers used the program intensely, with total investment through this scheme standing at a remarkable 1.5% of the GDP in the first five years of the policy. New units sold were typically high-quality flats in multi-family buildings marketed to mid-high/high-income households. We use this policy as an exogenous shifter in the spatial distribution of residential construction to induce exogenous variation in new housing around existing stores and supermarkets. This strategy gives us an instrument to estimate the effect of new local stock on retail prices, varieties and entry.[1]

We first test whether the introduction of new residential units influenced the price of groceries available locally to consumers. We find new stock leads to a reduction in grocery prices. Our instrumental variable estimates point to an elasticity of prices to new housing area of between -3% and -4%. Thus, our findings indicate that an increase in housing stock results in higher purchasing power for incumbent households in the vicinity of affected stores.

This result appears counter-intuitive, as local prices respond negatively to what we interpret as a positive demand shock. To rationalize this finding, we introduce a theoretical framework based on Mayer et al. (2014) in which multi-product firms competing in quantities face an increase in local demand. In our framework, this increase in demand can lead to a reduction in markups if there is either an increase in entry or an increase in the varieties available to consumers.

We then use our empirical strategy to test for these predictions. We find evidence of a large increase in available varieties at the local level in neighborhoods receiving a residential development shock. The elasticity of available product varieties to newly built area amounts to 17%. In terms of entry, we find a transitory effect on the number of grocery stores available in affected areas but this largely dissipates by the end of our sample period. The result of the entry and subsequent exit of stores during this period is a change in the composition of stores available locally: the average number of cash registers in nearby stores increases substantially after newly built properties come on the market.

Taken together, our results indicate that the local increase in demand induced by the change in housing stock improves the retail landscape for households in these neighborhoods: the price of groceries experience a moderate reduction, the varieties available increase substantially, and there is no change in the convenience of access. In light of this evidence, public concerns about the negative effect of neighborhood change on equitable access to groceries

[1] For an analysis of the specific effects of the policy on housing markets, see work in González-Pampillón (2021).

are not warranted.

Our analysis is carried out using a detailed product-level database of daily posted prices compiled by the General Directorate of Commerce (DGC, by its Spanish acronym), a branch of the Ministry of Economy and Finance in Uruguay. The data comprises detailed information from grocery stores all over the country, including hundreds of stores in Montevideo. An advantage of this database relative to the scanner data popular in most studies of retail markets and prices in developed countries is coverage: because the retail landscape in Montevideo includes a series of medium and small stores alongside larger supermarket chains, scanner data platforms have incomplete coverage in this context.

This paper contributes to the growing literature on urban consumption that emerged following the seminal contribution in Glaeser et al. (2001). Some strands of this literature focus on endogenous consumption amenities (Diamond, 2016; Guerrieri et al., 2013; Almagro and Domınguez-Iino, 2019). Allcott et al. (2019) use a structural model of grocery demand to conclude that differences in the supply of groceries at the local level only explain a small fraction of nutritional inequality in the United States. Closer to our paper, Handbury and Weinstein (2015) and Handbury (2019) study differences between US cities in both the prices of goods and the varieties available to consumers across the income distribution. Our paper distinguishes itself from this literature by looking specifically at the impact of neighborhood change on local grocery supply conditions, where neighborhood change is brought about by the physical transformation of neighborhoods by new residential development. The focus on Montevideo allows us to leverage credibly exogenous variation in the distribution of new building activities in estimation.

This paper is also related to the growing literature on the effect of gentrification and neighborhood change on local outcomes. Previous work in this literature has analyzed residential mobility patterns in gentrifying neighborhoods attempting to measure the extent of displacement of original residents. A group of studies finds little evidence of higher out-migration of these residents (Vigdor, 2002; Freeman, 2005; McKinnish et al., 2010; Ellen and O'Regan, 2011a,b; Ding and Hwang, 2016). Three recent studies (Aron-Dine and Bunten, 2019; Waights, 2018; Brummet and Reed, 2019) find instead that gentrification indeed leads to out-migration and displacement. Brummet and Reed (2019) also show that original home-owners who stay after the neighborhood gentrifies benefit from higher house values and increased employment levels. Autor et al. (2017) estimate the causal effect of gentrification induced by a rent deregulation policy on crime, finding a substantial reduction of crime rates. Closer to our work here, Asquith et al. (2021) study the effect of new residential stock on local housing prices and rents, finding a depression of local rents despite the new stock being occupied by relatively high-income residents. Our contribution to this broad literature is to look specifically into how neighborhood change affects local retail options for households.[2]

Finally, this paper is also related to previous work that estimates the effect of changes in (local) house prices on local retail prices. Stroebel and Vavra (2019) estimate how changes in house prices affect local retail prices. They argue that their estimates are not driven by changes in demographic or gentrification patterns, pointing to changes in the behavior of existing home-owner residents due to changes in their housing wealth given by changes in house prices, which lead firms to increase mark-ups in response. While we also look at interactions between housing and retail markets, we instead study how a process of physical change in neighborhoods affects retail prices in local stores.

2. Institutional Setting and Data

2.1. Institutional Setting

This section describes the place-based policy that underpins our strategy to study the effect of new housing stock on local grocery markets.

In August 2011, the Uruguayan government passed Law 18,795, entitled Ley de Acceso a la Vivienda de Interés Social (which roughly translates to Access to Housing of Social Interest Law, henceforth LVS for its Spanish Acronym).[3] The LVS aims at increasing the stock new build housing by means of a series of place-based tax benefits for the development of new residential units. Developers and private investors building new stock in certain locations are exempted from paying corporate tax (25% rate) on profits made on the sale of the new housing units, while house rents are partially exempted from personal income and corporate taxes for a period of 9 years.[4] Under the scheme, 540 new construction projects were promoted from December 2011 until December 2018, involving almost 17,000 new units. The total amount invested during this period rose to almost USD 1.4 billion, amounting to

[2]The link between retail access and neighborhood change has also been studied by the urban planning literature on *retail gentrification*. See for example, Mermet (2017), Zukin et al. (2009) and González and Waley (2013). These studies tend to focus on how the entry of boutique or gourmet shops replaces traditional retailers rather than on the effect of neighborhood change on the prices and varieties of grocery goods available locally. Despite this difference in focus, our finding that new development results in the entry of larger retailers is relevant for this line of work.

[3]The word *social* here is somewhat misleading. As discussed below, the vast majority of new units built under the aegis of the law were marketed to middle or middle-high income households.

[4]Other minor fiscal advantages include the exemption of the wealth tax over land and improvements during construction, as well as, over produced and subsequently rented units until nine years. They are also exempted to pay the transfer tax in case of buying unsold units. Finally, the law establishes tax credits for value-added tax on national and imported inputs.

roughly 1.5% Uruguayan GDP. The city of Montevideo concentrated 65% of the total projects (349 projects).[5]

The LVS policy can be used to subsidize projects of up to 100 new units by land lot. However, there are exceptions made for projects performed in large vacant lots or in parcels with abandoned housing or factories. Anecdotal evidence suggests many of the projects funded through the policy in the past decade indeed used parcels with vacant or derelict buildings. Eligibility conditions include unit size restrictions dependent on the number of bedrooms (i.e., between $32m^2$ and $50m^2$ for one bedroom units, increasing with each additional bedroom up to four).[6] LVS units also had to adhere to the guidelines laid down in the National Housing Plan and other ministerial regulations on quality. Compliance of LVS projects with these conditions was enforced via a vetting process involving the National Housing Agency (ANV) the Ministry of Economics and Finance and the Ministry of Housing. The resulting units created under the aegis of the scheme were usually high quality apartments in multi-family developments. Appendix Figure A.1 shows the distribution of quality for the LVS units and the existing stock in Montevideo. Around 95% of the LVS units were assessed as 'Excellent' by the Municipal Property Registry, while the average non-LVS dwelling for the city is assessed as having regular quality. The average time between the approval of a new project and the completion of building activities was of 21 months.

Eligibility for the subsidy for new construction offered by the LVS policy is place-based. The relevant regions in Montevideo are shown in Figure 1. The tax benefit only applies in the area labeled as S, which represents 52% of the total urbanized area, and is composed of both central and peripheral neighborhoods. This area is highly heterogeneous in income, with a coefficient of variation of 30% using per capita disposable household income. The unsubsidized area labeled as U in Figure 1 comprises most of the high-income neighborhoods in the city, with an average real per capita income that doubles the one in area S. Appendix Figure A.4 shows that this pattern is also observed for housing prices.[7] Figure 1 displays the spatial distribution of the LVS projects.

The boundaries of the subsidized area were defined jointly by the Ministry of Housing, the Ministry of Economics and Finance, and the Local Government of Montevideo's City Council. While there are no official documents explaining how the delimitation of the LVS borders was established, the border follows along some of the city's main avenues.

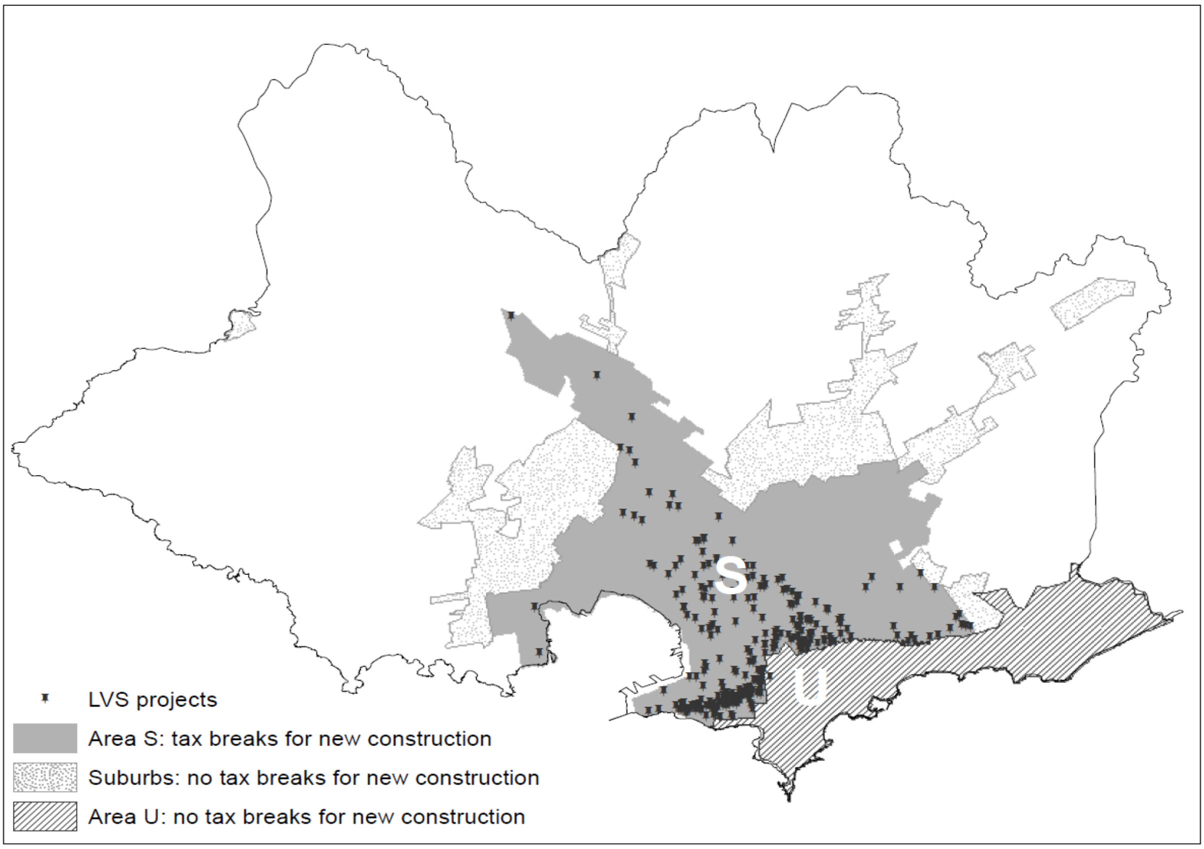
[5]Figure A.3 shows an example of a project performed in Montevideo before and after its implementation.

[6]Subsequent changes in regulation increased the lower bound of one-bedroom LVS units to $35m^2$.

[7]The unsubsidized area U also has better quality housing stock on average (see Appendix Figure A.2). Quality measured by the local municipal register based on structural property characteristics.

Figure 1

Place-based scheme for new construction projects in Montevideo (Uruguay)



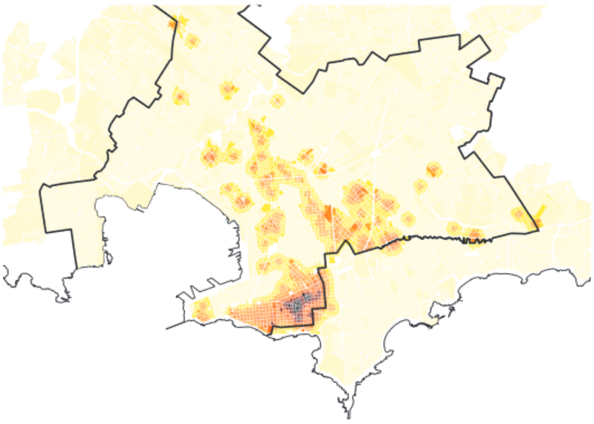
Notes: The policy was introduced in August of 2011. The subsidy for new construction projects only applies in the grey-area S. Development in area U received no exemptions. Black markers correspond to LVS projects approved for development in the period 2011-2018.

In what follows, we will use the LVS as a source of exogenous variation in the location of new residential development in Montevideo. Figure 2 illustrates how the LVS policy shifted new construction activity in the city. Panel A shows a heatmap for LVS projects carried out between the onset of the policy and 2018. We can observe these are located in the eligible region and, in most cases, are concentrated close to the region's boundary. Panel B illustrates how the policy induced a change in *overall* construction activity near the this boundary. For census tracts at different distance bands around the boundary, we calculate the average change in new residential area built between the pre-policy period (2004-2010) and the period in which LVS properties came on the market (2013-2019). We plot the change in the vertical axis against distance to the boundary, with positive distances corresponding to census tracts in the eligible region. We can observe that areas within the eligible region and close to the boundary experienced a significant increase in total new building activity in these periods. This is the spatial variation that we will use for identification.

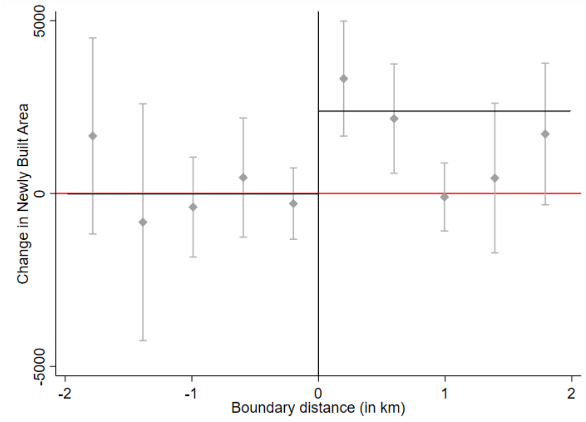
It is also important to highlight the temporal structure of the shock to local housing

Figure 2

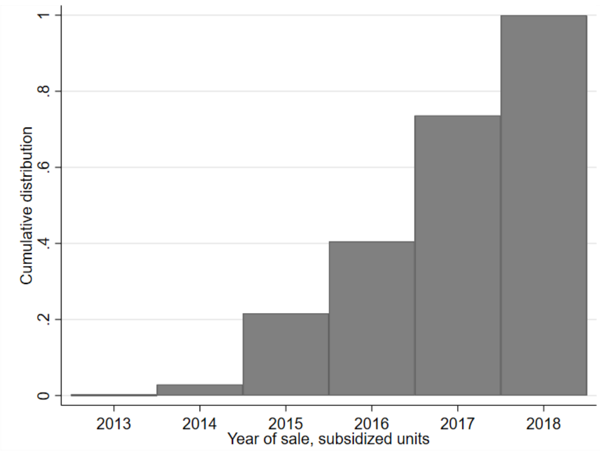
LVS and New Building Activity in Montevideo



(a) LVS Projects Heat Map



(b) Change in All New Building



(c) Timing of LVS Property Sales

Note: **Panel A** presents a heat-map of LVS projects in the city of Montevideo. The solid black lines corresponds to the boundary of the region eligible for LVS housing development. **Panel B** illustrates changes in construction activity between the 2004-2010 and the 2013-2019 periods, as measured using information from the municipal property registry. The horizontal axis represents distances to the LVS region boundary with negative distances corresponding to locations outside this region and positive distances to locations inside the regions. Black markers correspond to binned averages by distance. Vertical bars correspond to 95% CIs for those averages. Solid horizontal black lines correspond to averages calculated on each side of the boundary. **Panel C:** Timing of market sales of units from LVS projects that were approved for development between 2011 and 2014. Vertical axis represents frequencies relative to all sales up to 2018. Own calculations based on combining official data on LVS projects with data on housing transactions from the National Registry Office for the period 2011-2018.

supply induced by the policy. The LVS vetting process, the time required to obtain building permits from the city government, and the protracted build times usually associated to multi-family developments, meant it took several years before the first LVS properties came on the market. The timing of accumulated final sales of units in LVS developments approved between 2011 and 2014 are displayed in the Panel C of Figure 3. We can see that very few sales – less than 5% – had taken place before 2015, and roughly 60% of sales did not come until 2017. As a result, our empirical strategy will only provide suitable exogenous variation in the stock of new properties in the final years of our sample, a factor we will take into account when using this variation to estimate the effect of changes in stock on local retail conditions.

2.2. Data

Our main dataset is based on a detailed product-level database of daily posted prices compiled by The General Directorate of Commerce (DGC, by its Spanish acronym), a branch of the Ministry of Economy and Finance in Uruguay, which comprises information about grocery stores all over the country.[8] The DGC is the authority responsible for the enforcement of the Consumer Protection Law and requires retailers to report their daily prices once a month using an electronic survey.

The database has its origins issued Resolution Number 061/006 by the DGC, which mandates that grocery stores and supermarkets report their daily prices for a list of products if they meet the following two conditions: i) they sell more than 70% of the products listed, and ii) they either have more than four grocery stores under the same brand name or have more than three cashiers in a store. The information sent by each retailer is a sworn statement, and there are penalties for misreporting. The stated objective of the government with these measures is to ensure that prices posted on the website reflect the actual posted prices in the stores. In this regard, stores are free to set the prices they optimally choose, but they face a penalty if they try to misreport them to the DGC.

The grocery prices data includes daily prices from April 1st of 2007 to December 31st of 2019 for 154 products, most of them defined by Universal Product Code (UPC). This detailed information allows us to track the same good in stores across the country, avoiding measurement problems resulting from different products being compared (see the discussion in Atkin and Donaldson (2015)). The product markets for the goods included in the sample represent 15.6% of the CPI basket. Most items have been homogenized to make them comparable, and each supermarket must always report the same item. For example, the Coca Cola soft drink

[8]This is an updated database from Borraz et al. (2014) and Borraz et al. (2016).

is reported in its 1.5 liter non-returnable container variety by all stores. If this specific variety is not available at a store, then no price is reported. The data are then disseminated on a public website that allows consumers to check prices in different stores or cities and compute the cost of different baskets of goods across locations.[9]

The three best-selling brands are reported for each product market.[10] Initial products were selected after a survey to some of the largest supermarket chains in the year 2006. Between 2010 and 2011, the list of products was updated, including some additional markets and reviewing the top-selling brands for good categories. The 154 products in the current database represent more than 60 markets defined at the product category level (e.g., sunflower oil and corn oil are considered as different product markets, and the same is true for 000 wheat flour and 0000 wheat flour). For some products, the information provided in the database does not identify the goods at the UPC level; e.g., in the meat and bread markets, products do not have brands. As a consequence, we keep the 127 products that can be successfully traced as identical in different stores (out of a total of 154 products). The detailed list of the 127 matched goods with their UPC, and the share in the Consumer Price Index (CPI) can be found in A.1. A total of 54 products entered the database in 2010-2011. Therefore, we will conduct our main analysis of price effects using two samples: 1) Our *balanced sample of goods* including the 73 unique grocery products consistently present from 2007 to 2019, and 2) a larger sample of 127 unique grocery products including those included in the price database in 2010.

The original price data has incomplete temporal coverage in 2007 – the first waves were disseminated in April of that year – so we limit our sample to the period 2008–2019.[11] In order to avoid price changes due to temporary discounts and sales, we calculate the mode monthly price for each product (see Eichenbaum et al. (2011)) for each product.

For each grocery store, we have information on its exact location, given by its coordinates, and whether it belongs to a supermarket chain. Our analysis will focus on the city of Montevideo, the capital and largest city of Uruguay, with nearly forty percent of the country's population. There are a total of 249 grocery stores and supermarkets located in urban areas Montevideo in the database. See Borraz et al. (2014) for a complete description of the supermarket industry in Montevideo. In most of our analysis, we restrict attention to grocery stores located within 2km of the LVS boundary, which leaves us with a total of 155 individual

[9]See http://www.precios.uy/servicios/ciudadanos.html and Borraz et al. (2014) for a detailed description of the database.

[10]Exceptions are sugar, crackers, and cocoa, which have only two brands; and rice, which has up to six brands. Supermarket own brands are not included in the dataset.

[11]We will provide tests of the sensitivity of our results to this choice in section 5.

stores.[12]

We complement our data on product prices by store with data on individual LVS projects, register data on housing transactions taking place in Montevideo and data on the municipal cadastre on the stock residential units in the city. These data are used either for descriptive purposes or – in the case of the cadastre – to measure the year in which new units were built.

Descriptive features of the database are reported in Table 1. Descriptive statistics for annualized price changes measured over the period 2010-2019, at the level of individual good-store pairs are reported in Panel A. We report averages for the balanced sample of goods and the full sample of goods as well as for averages computed with and without CPI weights applied at the store level by product category (see section 3.1). Average annualized price changes calculated in this way vary between 7.8% and 9.5%. This is broadly consistent with annualized inflation between 2010 and 2019 that stood at 8.2%. Panel B displays descriptives for the number of varieties sold by each store in 2010 and 2019. Again, we report figures for the balanced sample and full sample of goods. Panel C provides figures for the changes in building activity taking place within 1km of every store in our sample. These are measured as percentage changes in the area built and the number of units between periods 2019-2013 and 2010-2004. We observe a positive average change in building activity which may be partly due to the influence of the LVS.

3. Empirical Analysis: Residential Development and Retail Prices

3.1. Empirical Strategy

The primary aim of our empirical analysis is to estimate the effect of demand changes resulting from new residential development on local grocery prices. By local grocery prices we mean the price of groceries sold by local stores.

One important identification problem when trying to detect this hypothesized causal link is that local demand for housing space in a given location can itself be affected by local retail prices and the mix of local varieties available to consumers. In addition, other confounders such as ease of transport access or crime levels may simultaneously affect housing demand and grocery supplies. In order to untie these knots, we exploit the change in the spatial distribution of new residential development induced by the LVS policy. To do this successfully. we will focus our attention on the 147 grocery stores located within a two kilometer band of the LVS boundary.[13] These areas are more comparable with each other than, for example,

[12]The number of stores varies by year somewhat due to entry and exit and opening of new branches.

[13]The band around the boundary considered in the analysis is illustrated in Appendix Figure A.5.

TABLE 1

DESCRIPTIVE STATISTICS – STORES WITHIN 2KM OF THE LVS BOUNDARY

		Mean	Median	Std. dev.
A. Annualized % Price Changes (2010-2019)				
Balanced Basket of Goods (Unweighted)		8.3	8.4	3.1
Balanced Basket of Goods (Store Weights)		9.5	10.1	2.6
All Goods (Unweighted)		7.8	8	2.6
All Goods (Store Weights)		8	8.1	2.6
B. Varieties by Supermaket				
Number of Varieties (Balanced Basket)	2010	37.1	38	5.7
	2019	35.5	35	8.2
Number of Varieties (All Goods)	2010	99.3	104	11.1
	2019	90	89	17.8
C. Change in Newbuilding Activity				
Δ in New Built Area <1km of stores (%)		213	27.7	941.2
Δ in New Built Units <1km of stores (%)		165.4	38.7	585.6
D. Other Dataset Characteristics				
Total Number of Supermarkets in Dataset	2010	112		
	2019	136		
Total Number of Goods in Dataset	2010	126		
	2019	122		

Notes: Descriptive statistics for the database on grocery good prices from the DGC. Panel A represents annualized growth rates in prices calculated between 2019 and 2010 for both the basket of goods present in the sample consistently since 2008, and including goods added during 2010. Panel B represents the average number of goods in each basket across supermarkets in 2010 and 2019. Panel C represents changes in building activity taking place within 1km of stores in the sample. Sample restricted to stores within 2km of the LVS boundary in all panels.

areas in the urban periphery. Moreover, as shown in Panel A of Figure 2, it is at this scale where our quasi-experimental variation can be leveraged for estimation.

In our main analysis, we study how changes in local housing supply around grocery stores affect prices of goods sold in those stores. We do this by using store-level time varying variables measuring the new building activity taking taking place within 1km of each store. The 1km band corresponds to roughly a 10-12 minute walk at a moderate pace. Structural estimates in Eizenberg et al. (2021) indicate retail demand is indeed local, with a 1km change in distance to a neighbourhood reducing retail demand from that neighbourhood by 35%.[14] We can illustrate how the LVS scheme affected changes in local supply around the stores by estimating the following event-study specification at the store-level:

$$Log(\text{New Units}_{st}) = \sum_{k=2008}^{2019} \rho_k Policy_s \times \mathbb{1}\{t = k\} + Policy_s + \delta_t + u_{st} \tag{1}$$

where $Log(\text{New Units}_{st})$ is the logarithm of the number of new units built in year t within

[14] We will evaluate the robustness of our main findings to this choice in section 5.

1km of grocery store s, as recorded in the municipal property register. Variable $Policy_s$ is an indicator taking value 1 if store s is inside the LVS region and 0 otherwise. Henceforth, we will call these the policy and comparison regions. Finally, δ_t represents time-effects for every year. The sum in the right-hand side of equation 1 includes a set of interactions between $Policy_s$ and year dummies. Therefore, coefficients ρ_k will measure the difference in building activity between the LVS eligible region and the comparison region relative to some benchmark period, which in our case will be 2010, the year before the introduction of the policy.

We illustrate estimates for the ρ_k coefficients and their corresponding 95% confidence intervals in Figure 3. Three important conclusions can be drawn from this figure. The first is that the difference in the relative intensity of construction activities around grocery stores on both sides of the LVS boundary was stable before the introduction of the LVS policy. That is, there are no apparent differences in the trends followed by the intensity of new residential development around stores in the policy and comparison regions. The second conclusion is that, by the end of the period, a large and persistent difference in the presence of new units has appeared, which is consistent with the descriptive evidence in section 2.1. This is the variation we will us to estimate our effects of interest. Finally, Figure 3 shows that difference in completions across locations appeared roughly 5 years after the introduction of LVS in 2011. This is again consistent with the evidence on LVS sales shown in section 2.1. It took more than 5 years for the incentives provided by LVS to translate into new market sales, largely because of the time required to produce new multi-family buildings.

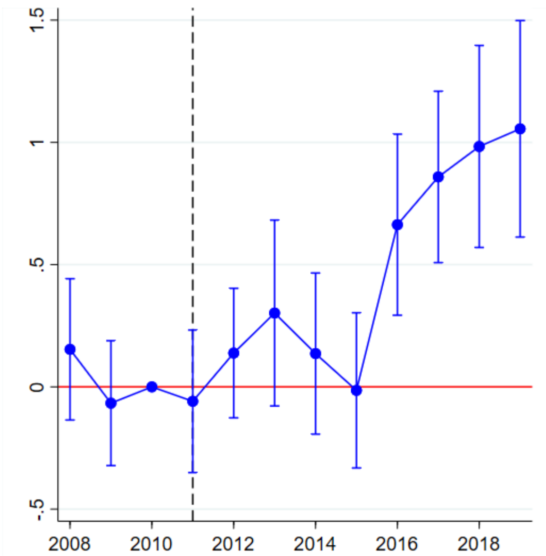
We now turn to discuss how we estimate the effect of new housing stock on local retail prices. In our first approach, we are interested in obtaining *reduced-form* estimates where we simply estimate the differences-in-differences of price levels in stores on different sides of the LVS boundary before and after the LVS properties came on the market. The estimating equation in this case can be written as:

$$Log(P_{ist}) = \beta_{RF} Policy_s \times post_t + \alpha Policy_s + \delta_{it} + \epsilon_{ist} \tag{2}$$

where P_{ist} is the price of product i in supermarket s and period t, $Policy_s$ is a dummy taking value one if supermarket s is located in the tax-exempt area, δ_{it} is a full set of product-time controls that accounts for aggregate product-type variation in prices, and coefficient α measures average relative price differences across locations before the policy was introduced. We will estimate this equation using data for 2010 – the year before the LVS policy was introduced – and 2019, the end date of our sample. These reduced-form estimates can be

Figure 3

Timing of New Residential Development



Note: Event-study coefficients from estimating equation 1. They measure the relative change in residential building activity (completions) within 1km of grocery stores between stores in the the LVS region and the comparison region for every year between 2008 and 2019. Effects are relative to 2010, the omitted year. Vertical segments correspond to 95% confidence bands. Dashed line corresponds to 2011 (the year the LVS was passed).

informative on the sign and significance of the effect of neighbourhood change on prices but they fall short of providing a quantitative estimate of the elasticity of prices to new housing stock.

Our second set of estimates address this issue by using the spatial and time variation in eligibility for the LVS tax-exemption to create an instrument for housing construction activity. Our instrumented variable New Area$_{st}$ measures the sum of the surface area (in m^2) of new units within 1km of supermarket s.[15] The variable is constructed using the accumulated stock of new units within six years of t (i.e., between $t-6$ and t).[16] We use the accumulated change over this period in an effort to measure changes to the density and vintage of the local housing *stock* rather than simply the *flow* change in construction in one given year. This variable measures the exposure of each supermarket s to new residential construction and, therefore, to changes in local demand for its goods. We estimate the effect of New Area$_{st}$ on local retail prices by estimating the parameter of interest β_{IV} via two-stage least squares (2SLS) where the two stages are given by:

[15]We can also measure new development using the *number* of new units built around each store. We will return to this alternative when discussing our robustness checks.

[16]We choose six years because the first new units built under the aegis of the LVS were sold in 2013, six years before 2019.

$$Log(\text{New Area}_{st}) = \pi Policy_s \times post_t + \eta Policy_s + \omega_{it} + u_{ist} \quad (3)$$

$$Log(P_{ist}) = \beta_{IV} Log(\text{New Area}_{st}) + \delta_{it} + \alpha Policy_s + \epsilon_{ist} \quad (4)$$

where equation 3 is the first-stage and 4 is the second-stage. Most variables in these equations are defined as above, with ω_{it} representing the product-time effects in the first stage. As with our reduced-form estimates, estimation is carried out using only the sample of stores within two kilometers of the LVS boundary. We include estimates of first-stage equation 3 in Appendix Table A.3 and report the associated F-statistics in our main tables.

It is straightforward to see that both for the reduced-form and IV estimates the identifying variation is the same: variation between regions across the boundary before and after the policy comes into effect. As a result, one of the identifying assumptions in both cases is a typical parallel trend assumption similar to the one in a conventional difference-in-differences study. We will show evidence in support of this assumption in the next section.

Regarding our estimates of price effects, we will report both unweighted and CPI-weighted estimates in the analysis. Weighting is important because the effective price faced by households buying a bundle of goods depends on the relative expenditure of each product in the household budget. As we do not observe household consumption at the individual level, we cannot compute these fractions directly or study changes in the share of income devoted to each product in response to the policy. What we do is use CPI weights of different product categories obtained from the Uruguayan National Statistical Office. As varieties of goods available may vary by supermarket and over time, we need to make suitable transformations to the original CPI weights if we want to use them to create appropriate baskets of goods.

We consider two alternatives. In the first place, we transform weights so that the total weight of a product category for a store or supermarket corresponds to the CPI weight of that category irrespective of the number of varieties sold in that store. This is quite straightforward and only requires re-scale these weights by the number of varieties available in each store at one point in time. For product a product belonging to product category k available in store s at time t we create $\omega_{kst}^{\text{store}} = \omega_k^{CPI}/n_{kst}$, where ω_k^{CPI} is the CPI weight for good category k and n_{kst} is the number of goods from category k present in our sample in period t and store s. We call these weights our *store* level weights because they vary both by product category and by store. Note that $\omega_{kst}^{\text{store}}$ ensures that the aggregate weight of all goods in a product category in a supermarket coincides with the weight of k in the CPI basket.[17]

An alternative is to build a weight that is fixed for every product category across all stores. We select this weight so as to ensure products that are relatively more widely available receive higher weights. We calculate $\omega_{it}^{\text{global}} = \omega_k^{CPI}\left(\frac{N_{it}^k}{I_{tk} \times N_t}\right)$ where N_{it}^k is the number of stores selling product i at time t, I_{tk} is the number of varieties in product category k and N_t is the number of stores open at time t. We call this our *global* weight because it is common for all product categories across stores. Note that if all goods are available in all stores at a point in time the global and store level weights will coincide.

3.2. Results

Before turning to our estimates of the effect of new developments on retail prices we first show graphical evidence in support of the parallel-trends assumption in our context. To do so, we estimate:

$$Log(P_{ist}) = \sum_{k=2008}^{2019} \phi_k Policy_s \times \mathbb{1}\{t = k\} + Policy_s + \delta_{it} + u_{ist} \tag{5}$$

where all variables are defined as above. The sequence of parameters $\{\phi_k\}_{t=2008}^{t=2019}$ captures the different paths of prices for stores on both sides of the LVS boundary relative to our reference year 2010. Estimates of these coefficients for the case without weights are reported graphically in Figure 4.[18] Panel A represents estimates obtained using the balanced sample of goods and Panel B represents estimates using the full sample of 127 goods. Both graphs show that the difference in grocery prices between stores in both regions around the LVS boundary was stable between 2008 and 2012. The p-value of a joint test for equality coefficients ρ_{2008} through ρ_{2012} is above 90% in both cases. This is reassuring as it indicates that the parallel trend assumption required for identification is satisfied in our context. We observe coefficients continue to be statistically insignificant in subsequent years up to 2016. As argued above, this is consistent with the fact that only a relatively small fraction of new LVS units had been effectively sold before 2017 – so that neither the local population density nor the composition of this population had been affected much by the policy yet. In 2017, we find a clear break from trend, with estimates shifting towards the larger reduced-form effects of around -3% that we observe in 2018 and 2019.

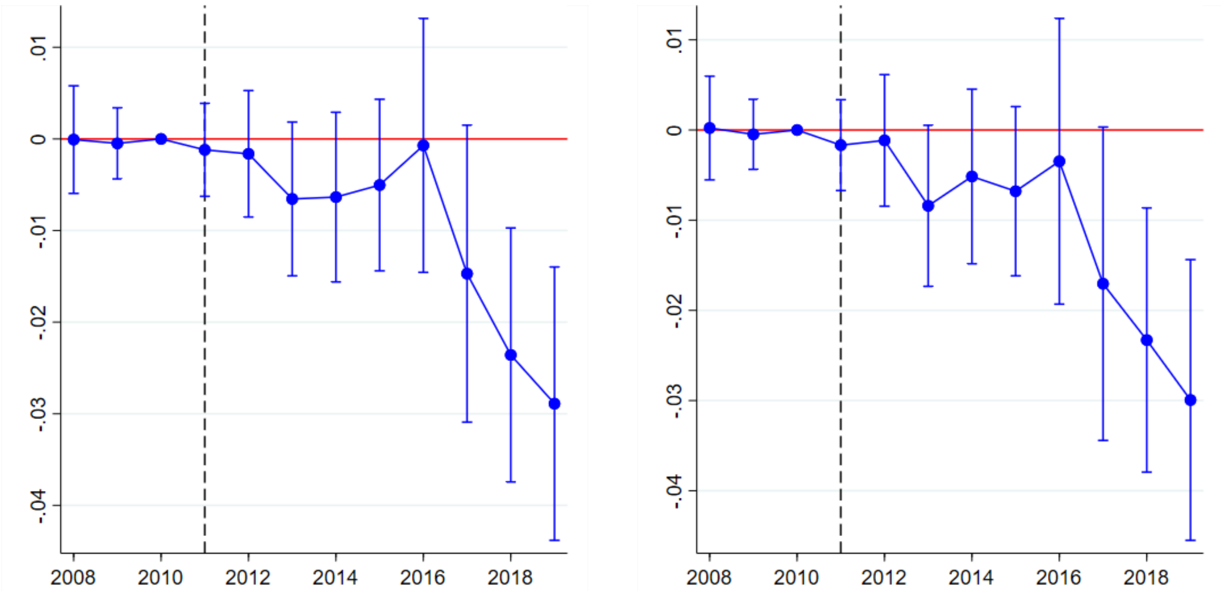
Quantitative estimates of the effect of new stock on prices are reported in Table 2. Columns

[17]To see this, note that $\sum_{i \in \Upsilon_{kst}} \omega_{kst}^{\text{store}} = n_{kst}\frac{\omega_k^{CPI}}{n_{kst}}$, where Υ_{kst} is the set of goods of category k present in our sample for store s at time t.

[18]Similar event-study graphs obtained using alternative specifications featuring store×product level weights, global weights and product-brand time effects are reported in section 5.

1 through 3 provide reduced-form estimates – see equation 2. Estimates reported in columns 2 and 3 are obtained using CPI store level product weights and global product weights respectively. We find *negative* and significant reduced-form effects on prices across the board, indicating that grocery stores located in the subsidized side of the LVS boundary reduced prices by roughly 2% relative to those on the comparison region.

FIGURE 4
EVENT-STUDY GRAPH: PRICES



(A) BALANCED SAMPLE OF GOODS　(B) FULL SAMPLE OF GOODS

Note: Reduced-form event-study type coefficients. Round markers indicate estimates for the sequence of ϕ coefficients in equation 5. Vertical bars correspond to 95% confidence intervals. Effects are relative to 2010, the omitted year. Vertical segments correspond to 95% confidence bands. Dashed lines corresponds to year 2011. Panel A represents estimates obtained with our sample of products consistently present in the sample from 2008. Panel B represents estimates obtained with the full sample of UPC-identifiable products.

Instrumental variable estimates of the elasticity of grocery prices to new residential development are reported in columns 4 to 6 of Table 2. Estimates in columns 5 and 6 obtained using store-level and global CPI based weights. The estimated elasticity of retail prices with respect new housing area ranges from -.031 to -.041.[19]

Taken together, these results confirm the findings illustrated in Figure 4. The new developments resulted in a moderate reduction of grocery prices available in nearby stores. This means that neighborhood change induced by the construction of new supply lead to moderate increases in purchasing power for incumbent households. Under the reasonable assumption – motivated by Engel's law – that households with relatively lower incomes spend more of

[19]Our first-stage estimates reported in panel A of Appendix Table A.3 indicate supermarkets in the policy region experienced an almost 60% increase in the area of new stock within 1km of their location relative to stores located in the comparison region. The instrument is reasonably strong, with an F-statistic of 21 in the specifications with and without weights.

those incomes in groceries, this effect can be positive for vertical equity across income groups. Thus, our results challenge the notion that neighborhood change will lead to a worsening of retail options to low-income incumbent households.

We discuss the mechanisms that could be leading to these findings in the next section.

TABLE 2
REDUCED-FORM AND IV ESTIMATES - GROCERY PRICE EFFECTS OF NEW DEVELOPMENTS

	Reduced-Form			IV		
	(1)	(2)	(3)	(4)	(5)	(6)
Policy $\times$ Post	-0.024***	-0.020**	-0.026***			
	(0.008)	(0.008)	(0.008)			
Log(New Area)				-0.038**	-0.031**	-0.041**
				(0.016)	(0.015)	(0.016)
CPI Weights	No	Store	Global	No	Store	Global
1st F-stat				21	21	21
Obs.	132192	132192	132192	132192	132192	132192

Notes: Standard errors are clustered at the store level. Estimates in columns 1 and 4 are obtained without using product weights. Estimates in columns 2 and 5 are obtained using store-level product weights based on CPI weights. Estimates in columns 3 and 6 are obtained using global product weights based on CPI weights. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

4. Mechanisms: Varieties and Entry

From the point of view of the local market for groceries, we interpret the construction of new housing stock induced by the LVS policy as an increase in local demand. Under this interpretation, our finding that grocery stores near the new developments reduced good prices appears counter-intuitive: a conventional supply and demand framework would make the opposite prediction in the face of an increase in demand.[20] Yet, this conventional framework may not be appropriate if other margins in the local supply of groceries can also respond (see e.g., Jaravel 2018).

These supply responses can operate via at least three channels: i) firm entry can lead to an increase in the number of stores, thus increasing competition and reducing markups as in a conventional Cournot model, ii) the introduction of new varieties within incumbent stores may change their pricing incentives and prompt a reduction in the price of previously available varieties, and iii) new housing development can affect the local price of land and thus reduce store's costs.

[20]This prediction is confirmed , for example, in recent work in Handbury and Moshary (2021) which reports a decline in grocery prices in response to a negative shock specific to breakfast and lunch product demand.

We explore the entry and variety mechanisms in what follows, first presenting a framework featuring endogenous prices, varieties and entry, and then analyzing these channels empirically by following a strategy similar to that used for price changes. Previous work in González-Pampillón (2021) suggests the land cost channel is unlikely to have played a role in this context: the policy itself, combined with local spillovers from new housing, increased local demand for land and lead to higher instead of lower prices of built-up space in the affected area.

4.1. Theoretical Framework

The trade literature on multi-product firms shows that an increase in market size can decrease prices, keeping the number of varieties constant (see Mayer et al. 2014). Separate work in the industrial organization literature (Ellickson, 2007) has shown that supermarkets increase the quality of the product offered when the market size increases. We draw on these intuitions when interpreting the price effects described in the previous sections as resulting from an increase in local demand for grocery stores' goods. This increase in local demand can arise via two channels. In the first place, the building of new multi-family units increases local densities and, therefore, the number of people living within existing store's local markets. In the second place, the fact that these units are new and generally of high quality (see section 2.1) implies they will attract relatively high income residents.[21] Evidence of positive spillover on housing prices are reported in González-Pampillón (2021).

To rationalize how an increase in local demand can result in lower prices, we propose a framework based on Mayer et al. (2014) where equilibrium prices are affected by number of offered varieties and entry of new competitors.[22] In our framework, changes in the scale of a market (i.e. the number of consumers available) result in lower prices via either of these channels.

There are L identical consumers with individual utility:

$$U = q_0 + \alpha \sum_j q_j - \frac{1}{2}\gamma \sum_j (q_j)^2 - \frac{1}{2}\eta \left(\sum_j q_j \right)^2 ,$$

where q_0 and q_j represents the individual consumption of the *numeraire* good and each variety j, respectively. The demand parameters α, γ, and η are all positive. Note that these preferences feature satiation points, i.e. utility becomes decreasing in q_j for large enough

[21] In our formal description below, we abstract from changes in local consumer types and simply treat this as a change in the scale of the market.

[22] A model using similar preferences has been recently used by Benkard et al. (2021) to explain the change in concentration in US product markets.

values of q_j. Maximizing utility we obtain the individual inverse demand for each variety:

$$p_j = \alpha - \gamma q_j^c - \eta Q. \tag{6}$$

where q_j^c is the individual consumption of good j and $Q = \sum_{i=1}^{N} q_i^c$, so the sum of individual consumption of all available varieties.

Production is carried out by identical firms that compete in quantities. In equilibrium, the relationship between individual consumption q_j^c and the supply by each firm q_j^m are given by $q_j^c = \frac{\sum_{k=1}^{M} q_j^k}{L}$, where M is the number of firms in this market. Substituting in individual demand, we obtain the demand function for each variety as a function of firm quantities q_j^k:

$$p_j = \alpha - \gamma \frac{\sum_{k=1}^{M} q_j^k}{L} - \eta \frac{\sum_{k=1}^{M} \sum_{j=1}^{N} q_j^k}{L} \tag{7}$$

Firms face entry costs F, fixed costs of offering each variety F_N and fixed marginal costs per unit c, with $c < \alpha$.[23] When considering the multi-firm equilibrium, we consider firms first entering simultaneously, then simultaneously choosing the varieties to be produced, and then simultaneously choosing quantities for each variety. Firm profits are therefore given by $\pi^m = \sum_{j=1}^{N_j} \left[q_j^m \left(p_j^m - c \right) \right] - F - F_N N$. Substituting the demand into the profit function, we can set up firm m's problem in the final stage (when choosing the quantity of each variety q_j^m:

$$\max_{\{q_j^m\}_{j=1}^{N}} \sum_{j=1}^{N} \left[q_j^m \left(\alpha - \gamma \frac{\sum_{k=1}^{M} q_j^k}{L} - \eta \frac{\sum_{k=1}^{M} \sum_{i=1}^{N} q_i^k}{L} - c \right) \right] - NF_v - F$$

Taking first-order conditions for this problem, we obtain:

$$\alpha - c - \frac{\gamma q_j^m}{L} - \frac{\gamma \sum_{k=1}^{M} q_j^k}{L} - \eta \left(\frac{q_j^m + \sum_{k=1}^{M} \sum_{i=1}^{N} q_i^k}{L} \right) = 0 \tag{8}$$

Solving for q_j^m we can obtain the reaction function for variety j sold by firm m. Note that the reaction function depends on the values of q_i^m for other varieties $i \neq j$. The specific functional form of this dependence derives from our choice of preferences, as do the results below.

We can use this framework to provide two comparative statics results, where we show how equilibrium prices, varieties or the number of firms vary with the number of consumers L. These are presented in Propositions 1 and 2.

[23] We can think of F_N as the fixed costs of sourcing and advertising each variety, and the cost of space associated to placing each variety at the store.

Proposition 1 - Market size, varieties and prices

Consider the problem of a monopolist choosing varieties and prices. In this case, a large enough increase in L results in an increase in endogenous varieties N and a reduction in the price of infra-marginal varieties.

Proof: See B.1.

The proof proceeds by obtaining an expression of firm profits as a function of varieties N. After characterizing the optimal number of varieties selected by the monopolist in this context N^*, we show this quantity increases with market size L (for sufficiently large changes in L). Finally, we show that this will result in a reduction in the markups for sold goods. Thus, we show that an expansion in the market for a retailer can lower prices via an expansion in varieties. It is worth noting that this mechanism relies on using preferences for which the product-level elasticity of demand increases (in absolute value) with the varieties of good available – i.e., additional varieties generate suitable substitutes for existing goods. We believe this is a reasonable assumption in the context of grocery markets.

Proposition 2 - Market size, entry and prices

Consider now the case in which the number of firms is endogenous. For a fixed number of varieties N, larger values of L result in more entry and lower equilibrium prices.

Proof: See B.2.

The proof proceeds by obtaining an expression for total firm profits as a function of the number of firms M. We characterize the equilibrium number of firms M^* and show that this figure is increasing in L. We also show that equilibrium prices are themselves decreasing in M^*, so that an increase in demand can lead to lower prices via its effects on entry, even if the number of varieties is fixed.

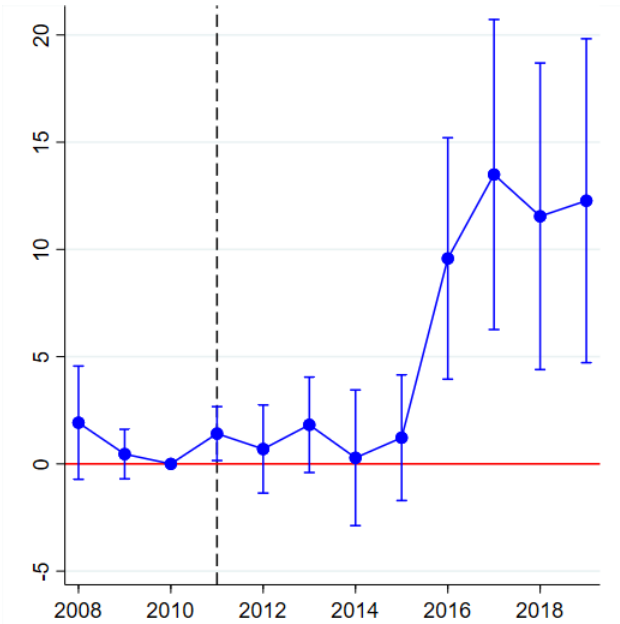
We have shown that both changes in varieties available or entry can provide scope for a reduction in prices resulting from a change in demand. Which of these mechanisms was behind our baseline results for the effect of new building activity in Montevideo? We turn to this question in the following sections.

4.2. Empirical Evidence: Change in Varieties

Informed by our theoretical framework, we now turn to test whether the introduction of new housing stock resulted in an increase in the varieties available to consumers locally. For this purpose we will exploit the same empirical strategy used in section 3, relying on exogenous variation induced by the shift in construction activity within the city. We measure varieties at the supermarket level, by calculating the percentage of reported products included in our price database that are offered at supermarket s and month t.

Before turning to our DID estimates for varieties, we report yearly coefficients akin to those reported in Figure 4, using our measure of varieties available as an outcome in a grocery store panel with interacted year effects. Coefficients for these interaction terms are illustrated in Figure 5, with effects being relative to 2010, the base year. As in the case of prices, we do not observe substantial changes in varieties available between both sides of the LVS boundary in the period between 2008 and 2012. We cannot reject the null that the coefficients for this period are equal to each other (p-value 36.1%). A substantial change is observed starting in 2016. Note that this coincides with the period in which we observe the break for new build sales. The coefficients for 2016 through 2019 are positive and large relative to those observed in the previous period, indicating an increase in varieties available for local consumers coinciding with the change in housing stock.

FIGURE 5
EVENT-STUDY GRAPH: VARIETIES



Note: Round markers indicate estimated coefficients from a regression of variety availability percentages at the store level on interaction terms between $Policy_s$ and year dummies featuring store and time effects. Effects are relative to 2010, the omitted year. Vertical segments correspond to 95% confidence bands.

To obtain the reduced-form and instrumental variable estimates of the effect of the change in housing stock on available varieties, we estimate the modified version of equations 2, 3 and 4 using a store panel for the years 2008 and 2019.[24] Estimates of the effect new residential

[24]For example, the reduced-form equation becomes

$$\text{Variety share}_{st} = \beta_{RF} Policy_s \times post_t + \delta_t + \alpha_s + \epsilon_{st}$$

.

development on the share of varieties offered by stores are reported in Table 3. Column 1 reports reduced-form estimates indicating that grocery stores in the side of the boundary that received the tax exemption for new development experienced a relative increase in varieties of roughly 12 percentage points. Column 2 reports IV estimates of the elasticity of the share of varieties available to new residential development. Results indicate that a one percent increase in newly built residential area within 1km of a store increases varieties available in that store by 0.17 percent.

We interpret these findings in light of the model presented in section 4.1. The change in housing stock prompted an increase in local demand for grocery stores, leading to an increase in varieties offered and a concomitant change in prices. Yet whether the increase in variety is the only mechanism explaining the change in prices requires exploring the role of entry. We turn to this in the next section.

TABLE 3
REDUCED-FORM AND IV ESTIMATES - PRODUCT VARIETIES AND NEW DEVELOPMENTS

	Reduced-Form	IV
	(1)	(2)
Policy $\times$ Post	12.395**	
	(4.806)	
Log(New Area)		17.167**
		(8.092)
First-stage F-stat		22
Obs.	232	232

Notes: Standard errors are clustered at the store level. First-stage F-statistic reported in column 2. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

4.3. Empirical Evidence: Entry

Changes in local housing stock prompt an increase in local grocery demand which can lead to the entry of new grocery stores in affected neighborhoods. The pro-competitive effects of entry may reduce local retail prices for residents, as shown in our theoretical framework. To investigate whether this mechanism explains our findings, we estimate the effect of the change in housing stock on access to grocery stores at the local level.

For this purpose, we compute two variables at the census tract level measuring the level of grocery store access in each year t. We first create variable Grocer Access$_{ct}^{1km}$ measuring the number of grocery stores open within 1km of the centroid of census tract c in year t. Alternatively, we consider variable

$$\text{Grocer Access}_{ct}^{1/d} = \sum_{s=1}^{S} \frac{D_{st}}{d_{cs}} \tag{9}$$

Grocer Access$_{ct}^{1/d}$ is an inverse-distance weighted average of access to grocery stores computed for each census tract c in every year t. S is equal to 249, the total number of stores in the urban areas of Montevideo, variable D_{st} is a dummy taking value 1 if grocery store s was active in year t, and d_{sc} is the Euclidean distance between store s and census tract c. Both Grocer Access$_{ct}^{1km}$ and Grocer Access$_{ct}^{1/d}$ are proxies for local access to grocery stores, with high values indicating access to a larger number of stores. Using both of these variable definitions and a census tract panel covering the period 2008-2019, we estimate our reduced-form equation:

$$Log(\text{Grocer Access}_{ct}) = \alpha_c + \delta_t + \beta_{RF} Policy_c \times post_t + \varepsilon_{ct} \tag{10}$$

where $Policy_c$ is a dummy taking value 1 if census tract c is located in the LVS policy region, α_c is a census tract fixed effect and δ_t represents year effect. The resulting estimate of β_{RF} will be positive if the number of grocery stores increases in areas affected by the LVS tax exemption. Estimates for this parameter for both of our outcomes are reported in columns 1 and 2 of Table 4. In addition, we report IV estimates of the effect of new residential development on entry, where new residential development is measured as the logarithm of the surface area of newly built stock in census tract c in the six years before year t.[25] The outcome variable is the log of the number of stores within 1km in columns 1 and 3 and the log of the inverse distance weighted access to grocery stores in columns 2 and 4.

Results in Table 4 lead us to conclude that the creation of new housing stock did not lead to a persistent increase in access to grocery stores. The number of stores available locally to households does not increase with new residential development and, by implication, there is no long-term effect in convenience. As discussed in section 5, these findings are robust to alternative ways of measuring new developments, changes in the baseline year and other methodological decisions. Yet this conclusion masks an interesting transitional pattern arising between 2011 and 2016. To explore how store availability changed over time during this period, we use our store access variables to estimate the event-study specification:

[25]Census tracts are relatively small geographies, with a total of 969 areas in the Montevideo, and over 450 areas within 2km of the LVS region boundary. In order to accommodate for the role of spatial dependence when conducting inference, we cluster at the level of $0.01^o \times 0.01^o$ cells. This leaves us with a total of 60 spatial clusters in the sample of census tracts within 2km of the LVS boundary.

TABLE 4
REDUCED-FORM ESTIMATES - GROCERY STORE ENTRY

	Reduced-Form		IV	
	(1) <1km	(2) 1/d	(3) <1km	(4) 1/d
Policy $\times$ Post	0.027	-0.008		
	(0.059)	(0.018)		
Log(New Area)			0.032	-0.005
			(0.074)	(0.023)
First-stage F-stat			48	48
Obs.	852	854	738	740

Notes: Estimates obtained from a census tract panel covering years 2010 and 2019. Standard errors are clustered at the level of $0.01^o \times 0.01^o$ grid cells. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

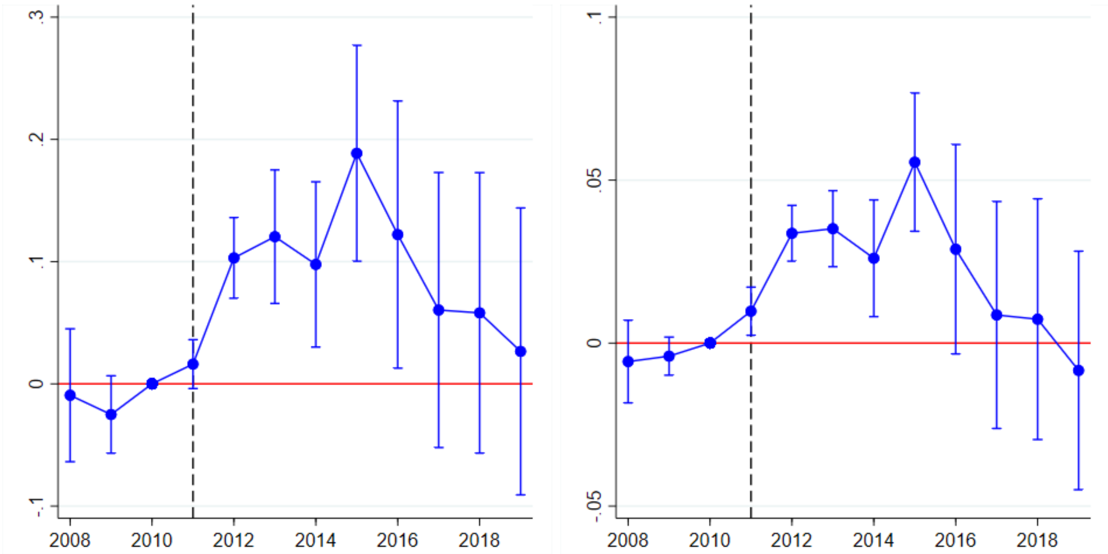
$$Log(\text{Grocer Access}_{ct}) = \sum_{k=2008}^{2019} \phi_k Policy_c \times \mathbb{1}\{t = k\} + \alpha_c + \delta_t + u_{ct} \tag{11}$$

Figure 6 plots the sequence of ϕ_t coefficients obtained when using the log of the number of stores within 1km (left panel) and the log of the inverse-distance weighted number of stores (right panel) as the outcomes in 11. We can observe that the introduction of the LVS policy did lead to a local increase in the number of stores initially, with access to grocery stores increasing after 2011 in the LVS region relative to the comparison region. Differential changes in access to stores peaks around 2015 and then drops, becoming not significant by 2019 in both graphs, in line with the results reported in Table 4.

We interpret this finding as suggesting that the (anticipated) change in housing stock led to a reshuffling of the types of grocery stores operating in the area. This dynamic aspect of the change in stores is not incorporated in our static theoretical framework but it suggests that the change in prices does not come (exclusively) from a change in the number of varieties offered by pre-existing stores but also from a response in the composition of stores available to consumers.[26] Hence, while the number of stores displays no long-run change, entry may have provided the adjustment margin for the change in varieties and the decline in prices to take place. To investigate this possibility, we conduct one additional exercise in which we estimate an event-study specification similar to the one in 11 but using as an outcome the average number of cash registers in stores within 1km of each census tract c in year t. This is a proxy for the size of grocery stores available locally.

[26]Interestingly, in their study of the US market, Glaeser et al. (2020) find evidence that gentrification increases the number of retail establishments, but it also triggers business closures.

FIGURE 6
EVENT-STUDY GRAPH: ENTRY



(A) STORES WITHIN 1KM (B) INVERSE DECAY DISTANCE ($1/d_{cs}$)

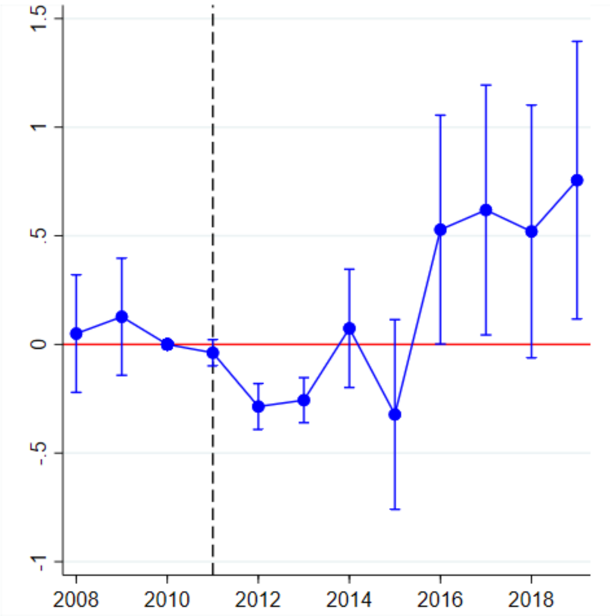
Note: Event-study graphs for changes in the number of stores available at the local level. Round markers indicate estimated coefficients from a census tract level regression of grocery shop access on interaction terms between $Policy_c$ and year dummies featuring census tracts and time effects (see equation 11). The effects displayed are relative to 2010 the omitted year. Vertical segments correspond to 95% confidence bands.

Figure 7 illustrates how the difference in the size of stores between the LVS and comparison areas changed over time. During much of the period between 2008 and 2015, this relative difference is fairly stable. However, in 2016 we see an abrupt increase in the average size of stores within 1km of census tracts located in the LVS region.[27] Note that this coincides in time with the sharp reduction in the estimates for local store density displayed in Figure 6. It also coincides with the increase in varieties reported in Figure 5. We interpret these findings as confirming that the entry and exit process taking place between 2010 and 2019 as a result of the increasing demand induced by new housing stock lead to a change in the local composition of stores and an increase in average store size.

Collectively, our empirical findings illustrate how neighborhood change induced by new housing stock shapes the local landscape of grocery supply. The increase in residential development activities leads to a period of store entry, followed by the exit of relatively small stores. As a result, there was a relative increase in long-term average store size at the local level. This increase in size led to an increase in locally available varieties and a concomitant

[27]Table A.6 in the Appendix reports both the reduced-form and IV estimates of the long-term effect of new development on grocery store size, which formally confirm the long-term changes in average store size.

Figure 7
Event-Study Graph: Number of Cash Registers in Nearby Stores



Note: Round markers indicate estimated coefficients from a average store size (number of cash registers) within 1km of a census tract on interaction terms between $Policy_c$ and year dummies featuring census tract and time effects. Effects are relative to 2010 the omitted year. Vertical segments correspond to 95% confidence bands.

reduction in grocery prices.

What drives the relationship between the change in varieties and the change in prices? We can consider two different mechanisms here. One is emphasized in the theoretical framework above: an increase in varieties can lead to pro-competitive effects because varieties are substitutes for each other and their increased availability increases the price elasticity of demand for each of them. Stores will respond by setting lower individual prices, thus reducing markups. An alternative mechanism that warrants attention would operate via lower costs of suppliers offering more varieties of goods. This can happen if, for example, larger stores have more market power when buying wholesale, something that is plausible in the case of grocery shops and supermarkets. To evaluate whether this mechanism is plausible in our case, we re-estimate the new development elasticity of grocery prices using the sample of stores that were consistently present between 2010 and 2019. This measures the pure pro-competitive effect resulting from the change in housing stock (Atkin et al., 2018). Under the assumption that the monopsony power of suppliers stays relatively constant over time, these estimates allow us to discriminate between explanations. In particular, if all of the price effect comes through an increase in presence of stores with market power in the wholesale market we should not observe a reduction in prices in continuing stores. Results reported in Table 5 show that the negative effect of prices discussed above is also observed in the sub-sample of continuing stores. We take this as evidence *against* the explanation of price effects based on upstream market power of suppliers.

TABLE 5
GROCERY PRICE EFFECTS OF NEW DEVELOPMENT – FIXING STORES

	Reduced-Form			IV		
	(1)	(2)	(3)	(4)	(5)	(6)
Policy × Post	-0.019**	-0.016**	-0.021***			
	(0.008)	(0.007)	(0.008)			
Log(New Area)				-0.026**	-0.021**	-0.028**
				(0.012)	(0.010)	(0.011)
CPI Weights	No	Store	Global	No	Store	Global
1st F-stat				27	33	30
Obs.	107374	107374	107374	107374	107374	107374

Notes: Estimation based on product-store-time level observations. Sample restricted to a fixed set of stores present in both 2010 and 2019. Estimates in columns 1 and 4 are obtained without using product weights. Estimates in columns 2 and 5 are obtained using store-level product weights based on CPI weights. Estimates in columns 3 and 6 are obtained using global product weights based on CPI weights. Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

5. Robustness Checks & Placebos

In this section, we provide a series of additional tests to evaluate the robustness of our findings. We will consider how our main results are affected by i) changes in the way we measure new building activity next to stores, ii) using an alternative baseline year, and iii) estimating price effects separately for low- and high-price brands. We also consider a series of placebo tests which rely on creating artificial areas obtained by shifting the location of the boundary in the eligibility areas of the LVS policy. Finally, we estimate the effect of new building on grocery prices keeping varieties available fixed.

Robustness Checks

We begin by considering the estimated effects of new development on grocery prices. Our baseline IV specifications in columns 3 and 4 of Table 2 use a definition of New Area$_{st}$ based on the sum of the m^2 of the new units within one kilometer of store s built in the 6 years prior to t. In Panel A of Table 6, we show that the 2SLS estimate for prices is robust to using the sum of the number of newly built units to measure quantities instead. The point estimates remain relatively close to those reported in our baseline results (see Table 2) and statistically significant at conventional levels.

In Panel B, we again use the sum of the m^2 of newly built units surrounding the grocery, but now change the time period to within five and seven years of period t – i.e., $t-5$ to t and $t-7$ to t. Once again, the resulting estimates do not differ compared to our baseline results, and our instrument still retains high-predictive power of new developments. In Panel C, we use an alternative definition of Log(New Area), using developments within 1.5 and .5 kilometers from each grocery s. We continue to find statistically significant reduction of retail prices in response to new development with elasticities between 2 and 4% across specifications. Finally, in Panel D, we use two alternative baseline years, 2009 and 2010. Estimates are still significant and magnitudes do not change considerably.

In tables A.4 and A.5, we repeat these four checks for our results on varieties and entry. In case of product varieties, estimates range from 12% to 21.3% compared to our baseline estimate of 12.4%, and being statistically significant at the 5% level in most the cases. The picture is similar in case of entry. In that case, all of the resulting estimates are not significant, confirming that the change in the number of stores between 2010 and 2019 was not concentrated in areas where new residential development was taking place.

We can use the data goods to explore whether the price effects documented above are concentrated on a particular subset of products within stores. In Appendix Table A.7, we estimate price effects separately for low- and high-price brands. As explained in Section 2.2, our database includes the three best-selling brands for each product market. We use variation in prices within product categories to define the high-price brand as the one with the highest average price across brands. Our definition of a high-price brand is likely to coincide with the definition of leader-brand. The remaining brands are then defined as low-price brands for exposition purposes. Results show a 3% reduction in the high-price brand and a similar decrease in low-price brands (3.6%), with point estimates not being statistically different from each other. This exercise shows that the overall price effects reported in Figure 3 are not driven by a particular type of product or market segment. Furthermore, these findings also have equity implications if households with different incomes consume products from different segments. As far as these issues are concerned, we do not observe substantial differences by segment.

Placebos

We can use the spatial nature of our empirical strategy to build a series of placebos. First, we construct a placebo border by shifting the original policy border southward until splitting the unsubsidised area U into two sub areas labelled as *Upper Placebo* and *Lower Placebo*. We can then use stores located in the unsubsidised area U, and we treat the *Upper*

TABLE 6

ROBUSTNESS CHECKS - PRICE EFFECTS

	Nbr. of Units			
	(1)	(2)		
A. New Units Instead of New Area				
Log(New Units)	-0.045**	-0.036*		
	(0.020)	(0.019)		
CPI Weights	N	Y		
First-stage F-stat	17	15		
Obs.	132192	132192		
	Time period: $[t-5,t]$		Time period: $[t-7,t]$	
	(1)	(2)	(3)	(4)
B. Time Period for New Stock				
Log(New Area)	-0.036**	-0.029**	-0.046**	-0.036*
	(0.015)	(0.014)	(0.020)	(0.019)
CPI Weights	N	Y	N	Y
First-stage F-stat	23	22	16	16
Obs.	132192	132192	132192	132192
	New housing within 1.5km		New housing within .5km	
C. Area Around Retail Store				
Log(New Area)	-0.041**	-0.034**	-0.025**	-0.020**
	(0.016)	(0.016)	(0.010)	(0.010)
CPI Weights	N	Y	N	Y
First-stage F-stat	31	25	22	21
Obs.	132192	132192	132192	132192
	Baseline Year: 2008		Baseline Year: 2009	
D. Alternative Baseline Year				
Log(New Area)	-0.042**	-0.029**	-0.042**	-0.029*
	(0.018)	(0.014)	(0.019)	(0.016)
CPI Weights	N	Y	N	Y
First-stage F-stat	16	18	18	20
Obs.	123029	123029	130395	130395

Notes: All estimates correspond to elasticities of prices to new development estimated via 2SLS. CPI weights included as indicated in each panel's foot correspond to product-store weights. Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 7
Placebo - Prices (Reduced-From Estimates)

	(1) Log(Price)	(2) Log(Price)	(3) Log(Price)	(4) Log(Price)
Post × Placebo	-0.001	0.003	-0.001	0.005
	(0.009)	(0.008)	(0.010)	(0.008)
Weights	N	Y	N	Y
Placebo	South	South	North	North
Obs.	60873	60873	42706	42706

Notes: Dependent variable is the logarithm of good prices measured at the store-month level. Columns 1 and 2 correspond to the placebo obtained by shifting the LVS boundary south. Columns 3 and 4 correspond to the placebo obtained by shifting the LVS boundary north. CPI weights included as indicated in each panel's foot correspond to product-store weights. Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Placebo area as the placebo policy region to test whether differences between these regions emerge in our outcomes of interest (see Figure A.9 for a graphical description). This first exercise is labeled as placebo *South* because that is the direction in which we displace the policy boundary. Results for retail prices are presented in columns 1 and 2 of Table 7, while results for varieties are presented in column 1 of Appendix Table A.8.

The second exercise – labeled as placebo *North* – is constructed by shifting the policy border northwards up to the centroid of the LVS subsidised area S (see Figure A.8 for a graphical description). In this case, we restrict our sample to stores within two kilometers of the artificial border which lie within the LVS area S. We build a binary variable that takes the value of one for stores located in the northern part of the placebo region and use this sample to test for differences in prices and varieties within regions. Results for prices of this placebo are reported in columns 3 and 4 of Table 7, and for varieties are reported in column 2 of Appendix Table A.8. All placebos yield statistically insignificant effects and point estimates that are substantially lower than those reported in our main analysis.

6. Conclusions

Neighborhoods are shaped by their physical characteristics, with an essential role played by housing in particular. Consequently, the introduction of new housing stock can induce a process of neighborhood change. Our results show that changes induced by large scale residential development activity affect the market for groceries faced by incumbent households. Specifically, we find evidence of a moderate *reduction* in grocery prices as a response to this change in demand induced by new housing development, as well as a substantial increase in available varieties for local residents.

Using our model, we show that these two facts can jointly arise in the context of a multi-product firm choosing what to produce: an increase in demand can prompt an expansion in the number of varieties offered and a reduction in prices. The model can be used to show that the reduction in prices can also result from the entry of new stores. While we find evidence of turnover in the participants in these local grocery markets, we do not find robust evidence of a sustained increase in the number of stores available locally as a result of the increase in housing stock.

The combination of a reduction in prices and an increase in varieties for fixed store density corresponds to a net improvement in the conditions for grocery consumers at the local level: Consumers can buy cheaper goods without a loss in the convenience of local access. Therefore, our results emphasize advantages of new development and neighborhood change for incumbent residents that have been largely overlooked by the literature. Moreover, they cast doubts on the risks that retail gentrification could pose for incumbent residents and their access to affordable groceries.

Our focus on conventional grocery goods – such as salt, soap, noodles, etc – implies that the changes in prices and varieties studied here will be especially relevant for low and middle-low income households for whom these goods amount to a larger share of their usual consumption basket. This makes our findings particularly relevant for the debate around the distributional consequences of neighborhood change. That being said, the fact that disaggregated spending data is not available in this context means we are unable to formally characterize the distributional impacts of these changes for different income groups. Efforts in this direction – which could follow recent developments in the study of inter-city differences in cost of living – remain an interesting avenue for future research.

Some final remarks are due regarding the external validity of our findings and, specifically, their *transportability* to other contexts (Pearl and Bareinboim, 2014). The use of the LVS policy as a source of exogenous variation yields clear advantages in terms of internal validity – it opens the space for a credible empirical strategy. The threats to external validity associated to this strategy are, as usual, less obvious. Most parameters of interest in this study are estimated off of variation in the development of multi-family buildings marketed to middle-high income households. Extrapolating our findings to the development of single-family neighborhoods or public/social housing may not be warranted. A different question is whether the mechanisms emphasized here can operate in general. The margins of adjustment of grocery supply will be available in most cities where land markets permit entry or changes in store size. These may be limited, however, in countries where urban planning systems impose tight restrictions along these margins (e.g. the United Kingdom (Cheshire et al., 2015)). Market structure may also be relevant. In Montevideo, the grocery market is characterized by the presence of three large supermarket chains and a large number of smaller players operating smaller stores (Borraz et al., 2016). Thus, the market structure in our context is comparable to that observed in other middle-sized cities in middle and high-income countries which retain a competitive fringe of independent stores. Keeping in mind these considerations, we remain optimistic about the replicability of our findings in other contexts. In any case, our results do show that new residential developments *can* improve access to groceries – in prices and varieties – to incumbent households.

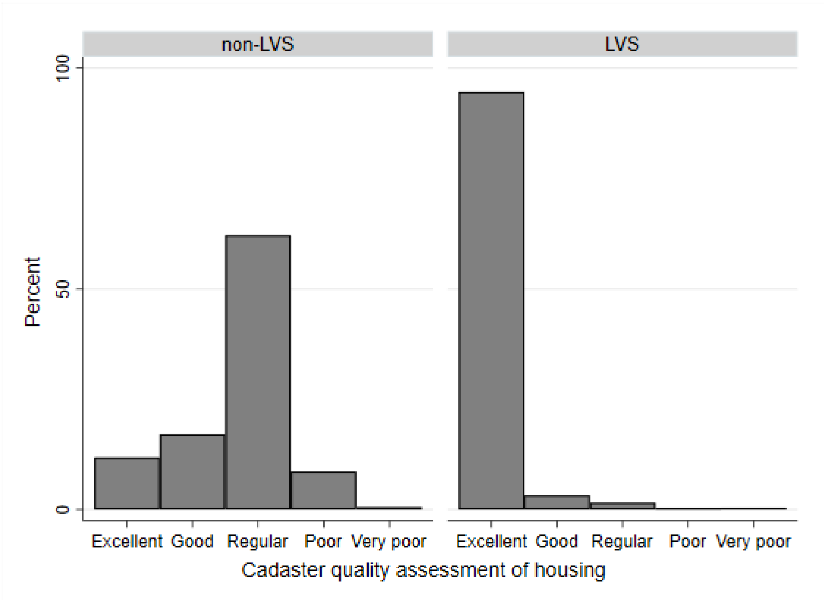
Online Appendices

A. Additional Figures and Tables

A.1. Quality of LVS units

FIGURE A.1

QUALITY OF LVS UNITS

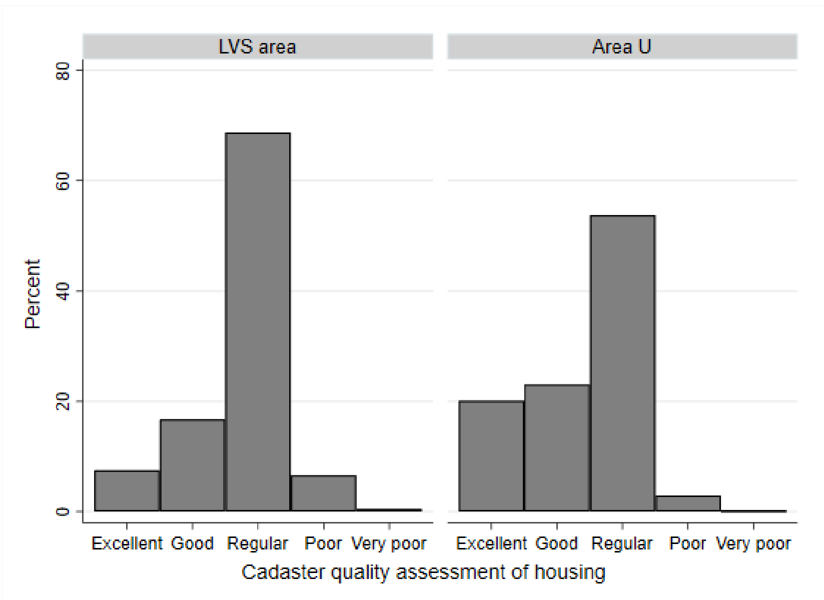


Source: Own calculations based on data from the Cadaster Agency (Municipal Property Registry).

Notes: The quality scale goes from 'Very poor' to 'Excellent'.

FIGURE A.2

QUALITY OF HOUSING WITHIN TWO KM OF BORDER $S - U$



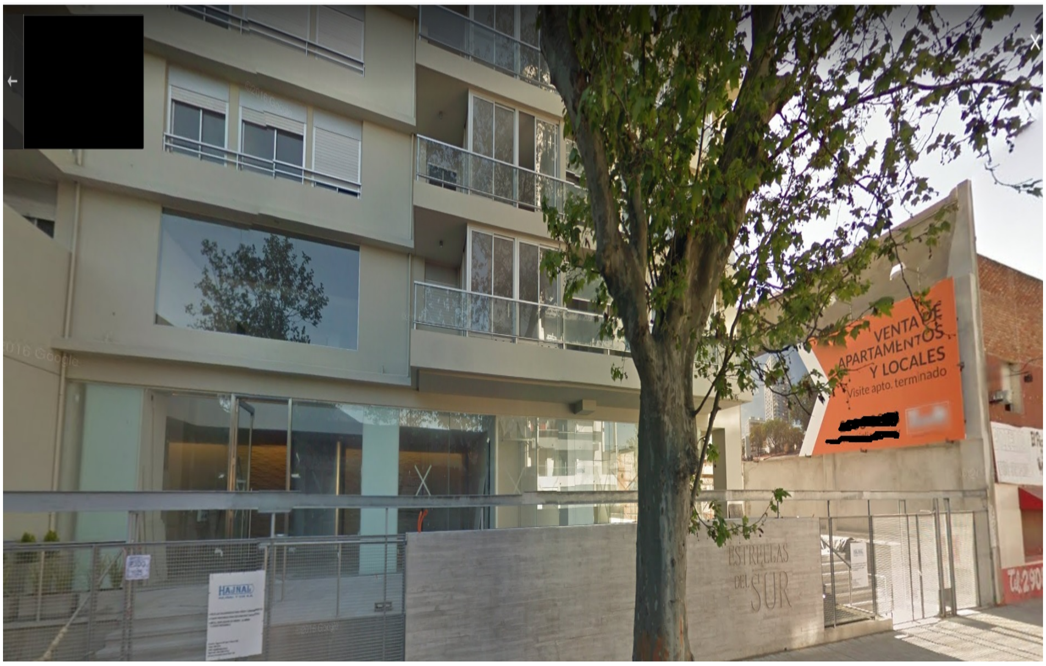
Source: Own calculations based on data from the Cadaster Agency (Municipal Property Registry).

Notes: The quality scale goes from 'Very poor' to 'Excellent'.

Figure A.3
Example of a LVS project

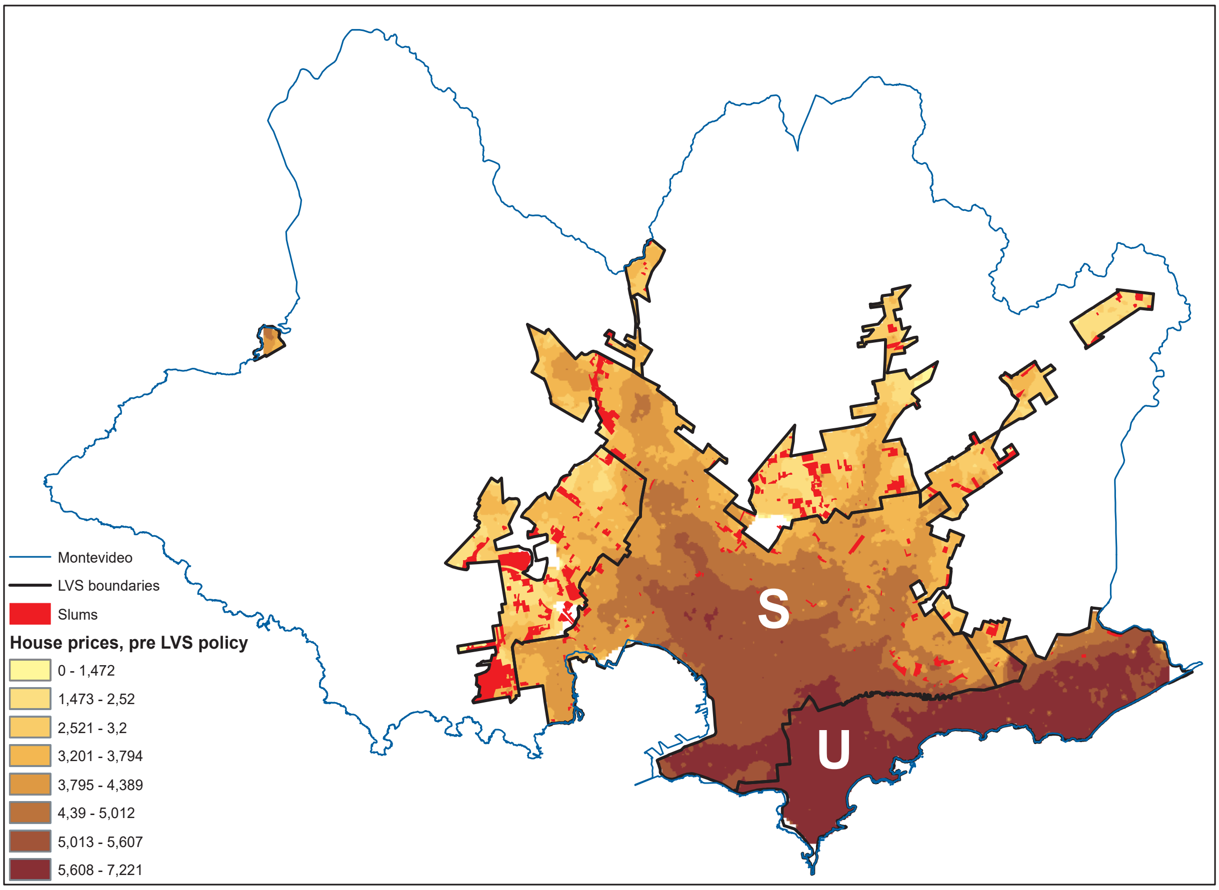
(a) Before

(b) After



A.2. House prices pre LVS policy

FIGURE A.4
MAP OF HOUSE PRICES (IN M^2, PRE LVS POLICY)



Notes: the map shows an inverse distance interpolation of the log of house prices (in m2) for the period 2004-2010, using grids of 100 *times* 100 metres and fixed search radius of 500 metres. Higher prices are represented with darker tones.

Table A.1
List of products

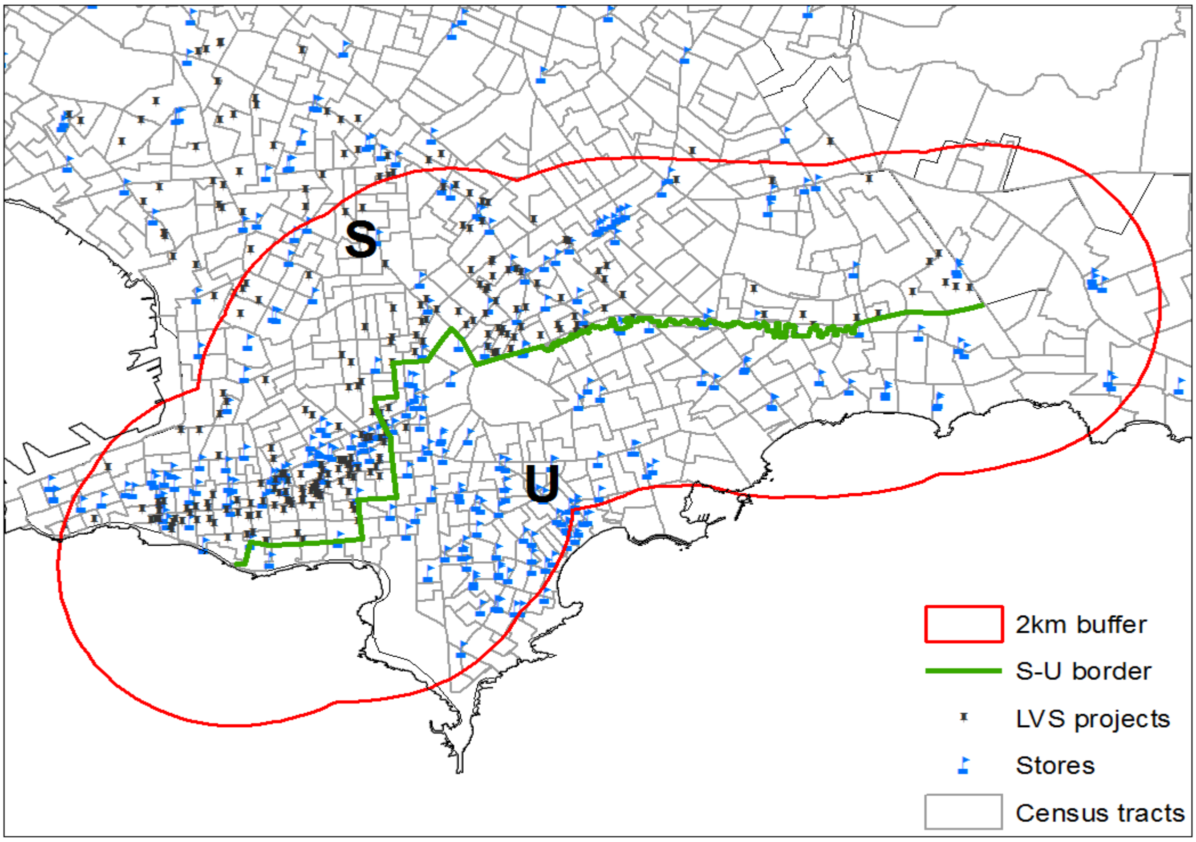
Product / Market	Brand	Specification*	UPC	% Share in CPI	Owner (/merger)	Sample Start (merge)
Beer	Patricia	0.96 L	7730452000435	0,36	FNC	2007/04
Beer	Pilsen	0.96 L	77302502	0,36	FNC	2007/04
Beer	Zillertal	1 L	7730452001319	0,36	FNC	2010/11
Wine	Faisán	1 L	7730540000187	0,80	Grupo Traversa	2007/04
Wine	Santa Teresa Clasico	1 L	7730135000035	0,80	Santa Teresa SA	2007/04
Wine	Tango	1 L	7730135000318	0,80	Almena	2007/04
Cola	Coca Cola	1.5 L	7730197232962	1,21	Coca Cola	2007/04
Cola	Nix	1.5 L	7730289000530	1,21	Milotur (CCU)	2007/04
Cola	Pepsi	1.5 L	7734284114087	1,21	Pepsi	2010/11
Cola	Coca Cola	2.25 L	7730197112967	1,21	Coca Cola	2010/11
Quince jelly	Los Nietitos	0.4 Kg	7730124020501	n/i	Los Nietitos	2009/01
Sparkling water	Matutina	2 L	7730922250070	0.81	Salus	2007/04
Sparkling water	Nativa	2 L	7730130000153	0.81	Milotur (CCU)	2007/04
Sparkling water	Salus	2.25 L	7730400000388	0.81	Salus	2007/04
Bread Loaf	Los Sorchantes	0.33 Kg	7730117000015	0,10	Bimbo / Los Sorchantes	2010/11
Bread Loaf	Bimbo	0.33 Kg	7730117001210	0,10	Bimbo	2010/11
Bread Loaf	Pan Catalán	0.33 Kg	7730230000336	0,10	Bimbo	2010/11
Brown eggs	Super Huevo	1/2 dozen	7730653000012	0,37	Super Huevo	2010/11
Brown eggs	El Jefe	1/2 dozen	7730637000045	0,37	El Jefe	2010/12
Brown eggs	Prodhin	1/2 dozen	7730239001211	0,37	Prodhin	2007/07
Butter	Calcar	0.2 Kg	7730901250176	0,22	Calcar	2007/04
Butter	Conaprole sin sal	0.2 Kg	77306197	0,22	Conaprole	2007/04
Butter	Kasdorf	0.2 Kg	7730105006357	0,22	Conaprole	2010/11
Cacao	Copacabana	0.5 Kg	7730109032154	0,07	Nestlé	2007/04
Cacao	Vascolet	0.5 Kg	7730109001686	0,07	Nestlé	2007/06
Coffee	Aguila	0.25 Kg	7730109012521	0,09	Nestlé	2007/04
Coffee	Chana	0.25 Kg	7730109012323	0,09	Nestlé	2007/04
Coffee	Saint	0.25 Kg	7730908360106	0,09	Saint Hnos	2010/11
Corn Oil	Delicia	0.9 L	7730132001196	n/i	Cousa	2010/11
Corn Oil	Río de la Plata	0.9 L	7730205040053	n/i	Soldo	2010/11
Corn Oil	Salad	0.9 L	7891080805738	n/i	Nidera	2010/11
Dulce de leche	Conaprole	1 Kg	7730105005091	0,13	Conaprole	2007/04
Dulce de leche	Los Nietitos	1 Kg	7730124384009	0,13	Los Nietitos	2007/04
Dulce de leche	Manjar	1 Kg	7730105005435	0,13	Manjar	2007/04
Flour (corn)	Gourmet	0.4 Kg	7730306000987	n/i	Deambrosi	2010/11
Flour (corn)	Presto Pronta Arcor	0.5 Kg	7790580660000	n/i	Arcor	2010/11
Flour (corn)	Puritas	0.45 Kg	7730354002322	n/i	Molino Puritas	2010/11
Flour 000 (wheat)	Cañuelas	1 Kg	7730376000085	0,16	Molino Cañuelas	2010/11
Flour 000 (wheat)	Cololó	1 Kg	7730213000506	0,16	Distribuidora San José	2010/11
Flour 0000 (wheat)	Cañuelas	1 Kg	7730376000061	0,16	Molino Cañuelas	2007/04
Flour 0000 (wheat)	Cololó	1 Kg	7730213000117	0,16	Distribuidora San José	2007/04
Flour 0000 (wheat)	Primor	1 Kg	7730133000105	0,16	Molino San José	2010/11
Grated cheese	Conaprole	0.08 Kg	7730105008832	0,14	Conaprole	2007/04
Grated cheese	Artesano	0.08 Kg	7730379000051	0,14	Artesano	2010/11
Grated cheese	Milky	0.08 Kg	7730153000185	0,14	Milky	2007/04
Deodorant	Axe Musk	0.105 Kg	7791293022130	0,27	Unilever	2010/11
Deodorant	Dove Original	0.113 Kg	7791293008141	0,27	Unilever	2010/11
Deodorant	Rexona Active Emotion	0.100 Kg	7791293004310	0,27	Unilever	2010/11
Hamburger	Burgy	0.2 Kg	7730138000575	n/i	Schneck	2010/11
Hamburger	Paty	0.2 Kg	7730901381146	n/i	Sadia Uruguay	2010/11
Hamburger	Schneck	0.2 Kg	7730138000599	n/i	Schneck	2010/11
Ice Cream	Conaprole	1 Kg	7730105912	0,24	Conaprole	2010/11
Ice Cream	Crufi	1 Kg	7730916580	0,24	Crufi	2010/11
Ice Cream	Gebetto	1 Kg	7730105980	0,24	Conaprole	2010/11
Margarine	Flor	0.2 Kg	7730132000571	n/i	Cousa	2010/11
Margarine	Doriana nueva	0.25 Kg	7805000300746	n/i	Unilever	2007/04
Margarine	Primor	0.25 Kg	7730132000533	n/i	Cousa	2007/04
Mayonnaise	Fanacoa	0.5 Kg	7790450086107	0,19	Unilever	2007/04
Mayonnaise	Hellmans	0.5 Kg	7794000401389	0,19	Unilever	2007/04
Mayonnaise	Uruguay	0.5 Kg	7730132000779	0,19	Unilever	2007/04
Noodles	Cololo	0.5 Kg	773021300	0,31	Distribuidora San José	2007/07
Noodles	Adria	0.5 Kg	773010330	0,31	La Nueva Cerro	2007/07
Noodles	Las Acacias	0.5 Kg	7730430000	0,31	Alimentos Las Acacias	2007/07

Table A.2
List of products (continued)

Product / Market	Brand	Specification*	UPC	% Share in CPI	Owner (/merger)	Sample Start (merge)
Peach jam	Dulciora	0.5 Kg	7790580508104	n/i	Arcor	2007/04
Peach jam	El Hogar	0.5 Kg	7730180086831	n/i	Lifibel SA	2010/11
Peach jam	Los Nietitos	0.5 Kg	7730124010304	n/i	Los Nietitos	2007/04
Peas	Campero	0.3 Kg	7730905130047	0,08	Regional Sur	2010/11
Peas	Cololó	0.3 Kg	7730213000018	0,08	Distribuidora San José	2010/11
Peas	Nidemar	0.3 Kg	7730332000975	0,08	Nidera	2010/11
Rice	Aruba tipo Patna	1 Kg	7730115170109	0,27	Saman	2007/04
Rice	Blue Patna	1 Kg	7730114000117	0,27	Coopar	2007/04
Rice	Green Chef	1 Kg	7730114400016	0,27	Coopar	2007/04
Rice	Pony	1 Kg	7730115020107	0,27	Saman	2010/11
Rice	Vidarroz	1 Kg	7730114000728	0,27	Coopar	2008/05
Rice	Saman Blanco	1 Kg	7730115040105	0,27	Saman	2010/11
Crackers	Famosa	0.14 Kg	7622300226480	0,25	Mondelez	2007/04
Crackers	Maestro Cubano	0.12 Kg	7730154000986	0,25	Bimbo	2007/04
Salt	Sek	0.5 Kg	77300607	0,08	Deambrosi	2007/04
Salt	Torrevieja	0.5 Kg	7730901390063	0,08	Torrevieja	2007/04
Salt	Urusal	0.5 Kg	7730214000062	0,08	UruSal	2007/04
Semolina pasta	Adria	0.5 Kg	77301030	0,31	La Nueva Cerro	2007/07
Semolina pasta	Las Acacias	0.5 Kg	7730430001	0,31	Alimentos Las Acacias	2007/07
Semolina pasta	Puritas	0.5 Kg	7730354001158	0,31	Molino Puritas	2010/11
Soybean oil	Condesa	0.9 L	7730132000434	0,09	Cousa	2008/05
Soybean oil	Río de la Plata	0.9 L	7730205067593	0,09	Soldo	2010/11
Soybean oil	Salad	0.9 L	7891080801693	0,09	Nidera	2010/11
Sugar	Azucarlito	1 Kg	7730251000018	0,24	Azucarlito	2007/04
Sugar	Bella Union	1 Kg	7730106005113	0,24	Bella Unión	2007/04
Sunflower oil	Optimo	0.9 L	7730132001165	0,29	Cousa	2007/04
Sunflower oil	Uruguay	0.9 L	7730132000441	0,29	Cousa	2007/04
Sunflower oil	Río de la Plata	0.9 L	7730205067661	0,29	Soldo	2010/11
Tea	Hornimans	Box (10 units)	7730261000046	0,08	José Aldao	2007/04
Tea	La Virginia	Box (10 units)	7790150572290	0,08	La Virginia	2007/04
Tea	President	Box (10 units)	7730220030527	0,08	Carrau	2010/11
Tomato paste	Conaprole	1 L	7730105015403	0,16	Conaprole	2007/04
Tomato paste	De Ley	1 L	7730306000604	0,16	Deambrosi	2007/04
Tomato paste	Gourmet	1 L	7730306000017	0,16	Deambrosi	2010/11
Yerba	Canarias	1 Kg	7730241003654	0,46	Canarias	2007/04
Yerba	Del Cebador	1 Kg	7730354000519	0,46	Molino Puritas	2007/06
Yerba	Baldo	1 Kg	7730241003920	0,46	Canarias	2010/11
Yogurt	Conaprole	0.5 Kg	7730105032820	0,13	Conaprole	2010/11
Yogurt	Parmalat (Skim)	0.5 Kg	7730112088520	0,13	Parmalat	2010/11
Yogurt	Calcar (Skim)	0.5 Kg	7730901250565	0,13	Calcar	2010/11
Bleach	Agua Jane	1 L	7731024003038	0,13	Electroquímica	2007/04
Bleach	Sello Rojo	1 L	7730494001001	0,13	Electroquímica	2007/04
Bleach	Solucion Cristal	1 L	7730377066028	0,13	Vessena SA	2007/04
Dishwashing detergent	Deterjane	1.25 L	7731024008118	0,11	Clorox Company	2007/04
Dishwashing detergent	Hurra Nevex Limon	1.25 L	7730165317424	0,11	Unilever	2007/04
Dishwashing detergent	Protergente	1.25 L	7730329024014	0,11	Electroquímica	2010/11
Laundry soap	Drive	0.8 Kg	779129078	0,35	Unilever	2007/04
Laundry soap	Nevex	0.8 Kg	779129020	0,35	Unilever	2007/04
Laundry soap	Skip, Paquete azul	0.8 Kg	77912902034	0,35	Unilever	2007/04
Laundry soap, in bar	Bull Dog	0.3 Kg (1 unit)	7791290677951	n/i	Unilever	2007/04
Laundry soap, in bar	Nevex	0.2 Kg (1 unit)	7791290677944	n/i	Unilever	2007/04
Laundry soap, in bar	Primor	0.2 Kg (1 unit)	7730205066	n/i	Soldo	2010/11
Shampoo	Fructis	0.35 L	78049600	0,31	Garnier	2007/04
Shampoo	Sedal	0.35 L	779129301	0,31	Unilever	2007/04
Shampoo	Suave	0.93 L	77912930083XX	0,31	Unilever	2007/04
Soap	Astral	0.125 Kg	7891024176771	0,14	Colgate	2010/11
Soap	Palmolive	0.125 Kg	7891024177XXX	0,14	Colgate	2007/04
Soap	Rexona	0.125 Kg	779129352XXXX	0,14	Unilever	2012/12
Toilet paper	Higienol Export	4 units (25 M each)	7730219001101	0,23	Ipusa	2007/04
Toilet paper	Elite	4 units (25 M each)	7790250021438	0,23	Ipusa	2010/11
Toilet paper	Sin Fin	4 units (25 M each)	7730219000494	0,23	Ipusa	2007/04
Toothpaste	Pico Jenner	0.09 Kg	7730366000170	0,17	Abarly / Colgate	2010/11
Toothpaste	Colgate Herbal	0.09 Kg	7891024133668	0,17	Colgate	2010/11
Toothpaste	Kolynos	0.09 Kg	7793100120121	0,17	Colgate	2010/11

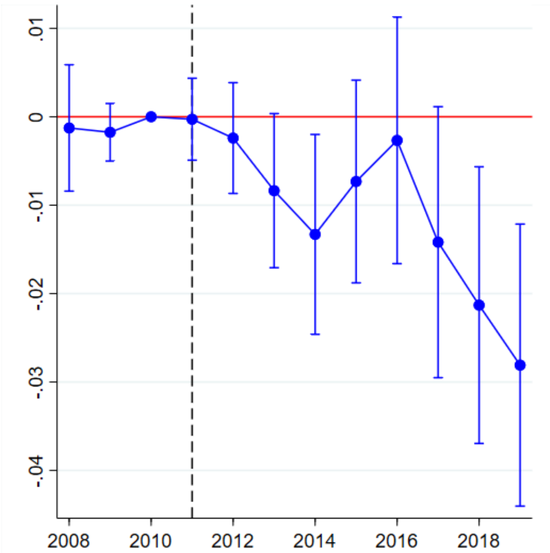
Kg = kilograms; L = liters; M = meters. n/i - No information.

Figure A.5
Area of the analysis

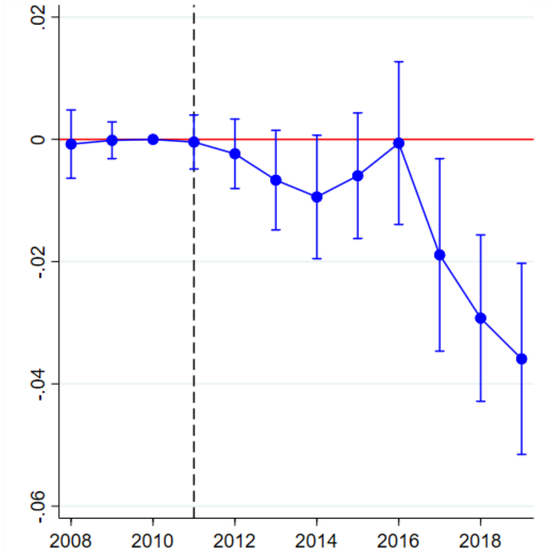


Notes: the area of the analysis is denoted by the 2km buffer (the red line). Then, units within this buffer are considered for the empirical analysis.

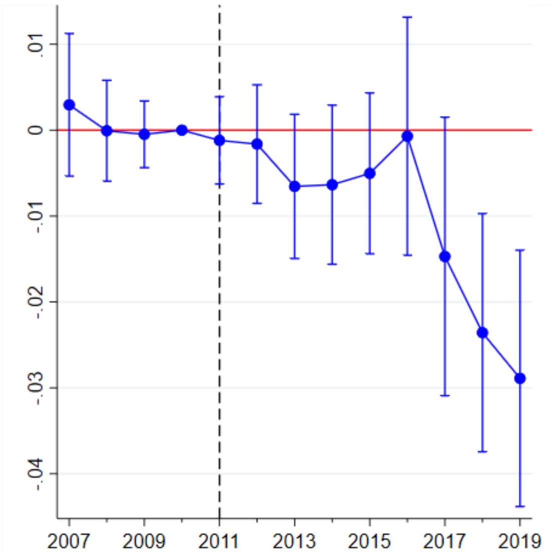
FIGURE A.6
EVENT-STUDY GRAPH: PRICES



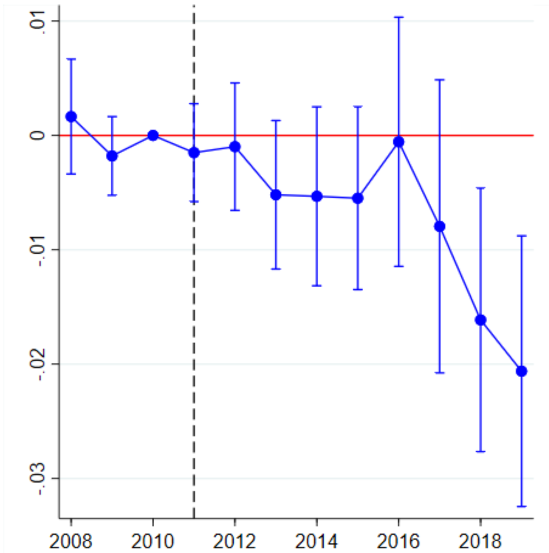
(A) USING STORE-LEVEL PRODUCT WEIGHTS



(B) USING GLOBAL PRODUCT WEIGHTS



(C) INCLUDING 2007



(D) PRODUCT-BRAND FIXED EFFECTS

Note: Reduced-form event-study type coefficients. Round markers indicate estimates for the sequence of ϕ coefficients in equation 5. Vertical bars correspond to 95% confidence intervals. Effects are relative to 2010, the omitted year. Vertical segments correspond to 95% confidence bands. Dashed lines corresponds to year 2011. **Panel A** represents estimates obtained using store-level product weights. **Panel B** represents estimates obtained using store-level product weights. **Panel C** represents estimates obtained after extending the sample from 2007 (incomplete year). **Panel D** represents estimates obtained in a specification featuring product-brand specific time effects instead of product group-time effects.

Table A.3

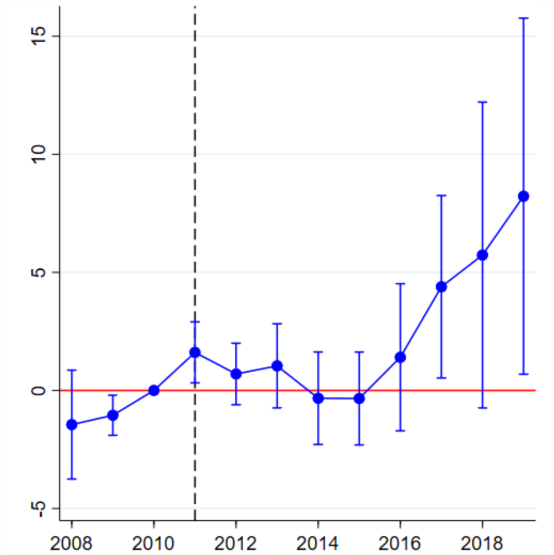
First-Stage - New Developments Effects of the LVS policy

	(1) Log(New Area)	(2) Log(New Units)
A. Product × month × store level		
Policy × Post	0.630***	0.715***
	(0.139)	(0.085)
First-stage F-stat	20	72
Obs.	132192	132192
B. Store × Year level		
Policy × Post	0.807***	0.684***
	(0.142)	(0.121)
First-stage F-stat	32	32
Obs.	170	170

Notes: Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Figure A.7

Event-Study Graph: Varieties. Fixed number of stores.



Note: Round markers indicate estimated coefficients from a regression of variety shares on interaction terms between $Policy_s$ and year dummies featuring store and time effects. Effects are relative to 2010 the omitted year. Vertical segments correspond to 95% confidence bands.

Table A.4

Robustness Checks - Product Varieties

	Nbr. of Units	
A. New Units Instead of New Area		
Log(New Units)	24.434*	
	(13.538)	
First-stage F-stat	18	
Obs.	225	
	Time period: $[t-5, t]$	Time period: $[t-7, t]$
B. Time Period for New Stock		
Log(New Area)	16.411**	20.344**
	(7.575)	(9.785)
First-stage F-stat	24	18
Obs.	232	232
	New housing within 1.5km	New housing within .5km
C. Area Around Retail Store		
Log(New Area)	19.837**	12.084**
	(8.826)	(5.613)
First-stage F-stat	30	24
Obs.	232	232
	Baseline Year: 2008	Baseline Year: 2009
D. Alternative Baseline Year		
Log(New Area)	19.990*	21.292**
	(10.333)	(9.982)
First-stage F-stat	16	20
Obs.	225	230

Notes: Store-level specifications using data for 2010 and 2019. Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

TABLE A.5

ROBUSTNESS CHECKS - ENTRY

	Nbr. of Units			
	(1)	(2)		
	<1km	1/d		
A. New Units Instead of New Area				
Log(New Units)	0.037	-0.006		
	(0.084)	(0.027)		
First-stage F-stat	30	30		
Obs.	738	740		
	Time period: $[t-5,t]$		Time period: $[t-7,t]$	
	(1)	(2)	(3)	(4)
	<1km	1/d	<1km	1/d
B. Time Period for New Stock				
Log(New Area)	0.025	-0.002	0.029	-0.005
	(0.065)	(0.021)	(0.081)	(0.026)
First-stage F-stat	47	47	47	47
Obs.	720	722	766	768
	Baseline Year: 2008		Baseline Year: 2009	
	<1km	1/d	<1km	1/d
C. Alternative Baseline Year				
Log(New Area)	0.013	-0.007	0.046	-0.005
	(0.082)	(0.025)	(0.073)	(0.022)
First-stage F-stat	45	45	46	46
Obs.	740	742	742	744

Notes: Census tract level specifications using data for 2010 and 2019. Standard errors are clustered at the census area level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

TABLE A.7

PRICE EFFECTS – HETEROGENEITY BY PRODUCT SEGMENT

	High-price brand		Low-price brand	
Log(New Area)	-0.030**	-0.022*	-0.036**	-0.024*
	(0.013)	(0.012)	(0.015)	(0.013)
CPI Weights	N	Y	N	Y
First-stage F-stat	21	18	20	21
Obs.	61293	61293	70887	70887

Notes: Instrumental variable estimates of the elasticity of grocery prices to new residential development. Subsamples of high-price (top priced) and low-price (other) goods for each product category as described in the main text. CPI weights in columns 2 and 4 correspond to product-store weights. Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

TABLE A.6

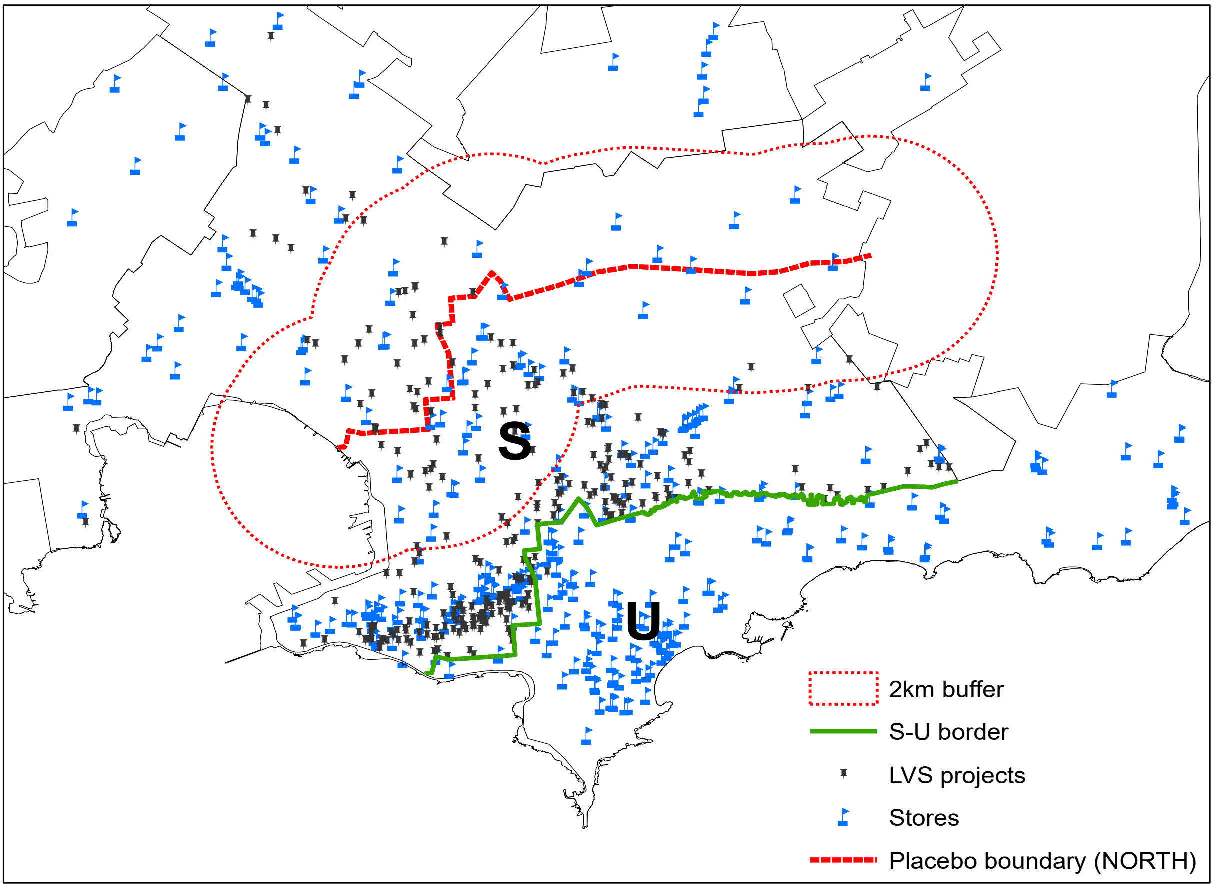
REDUCED-FORM & IV ESTIMATES - GROCERY STORE SIZE

	Reduced-Form	IV
	(1)	(2)
Policy × Post	0.756**	
	(0.319)	
Log(New Area)		0.916**
		(0.447)
First-stage F-stat		48
Obs.	852	738

Notes: Estimates obtained from a census tract panel covering years 2010 and 2019. The dependent variable is the average size of stores within 1km of census tracts. Standard errors are clustered at the level of $0.01^o \times 0.01^o$ grid cells. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

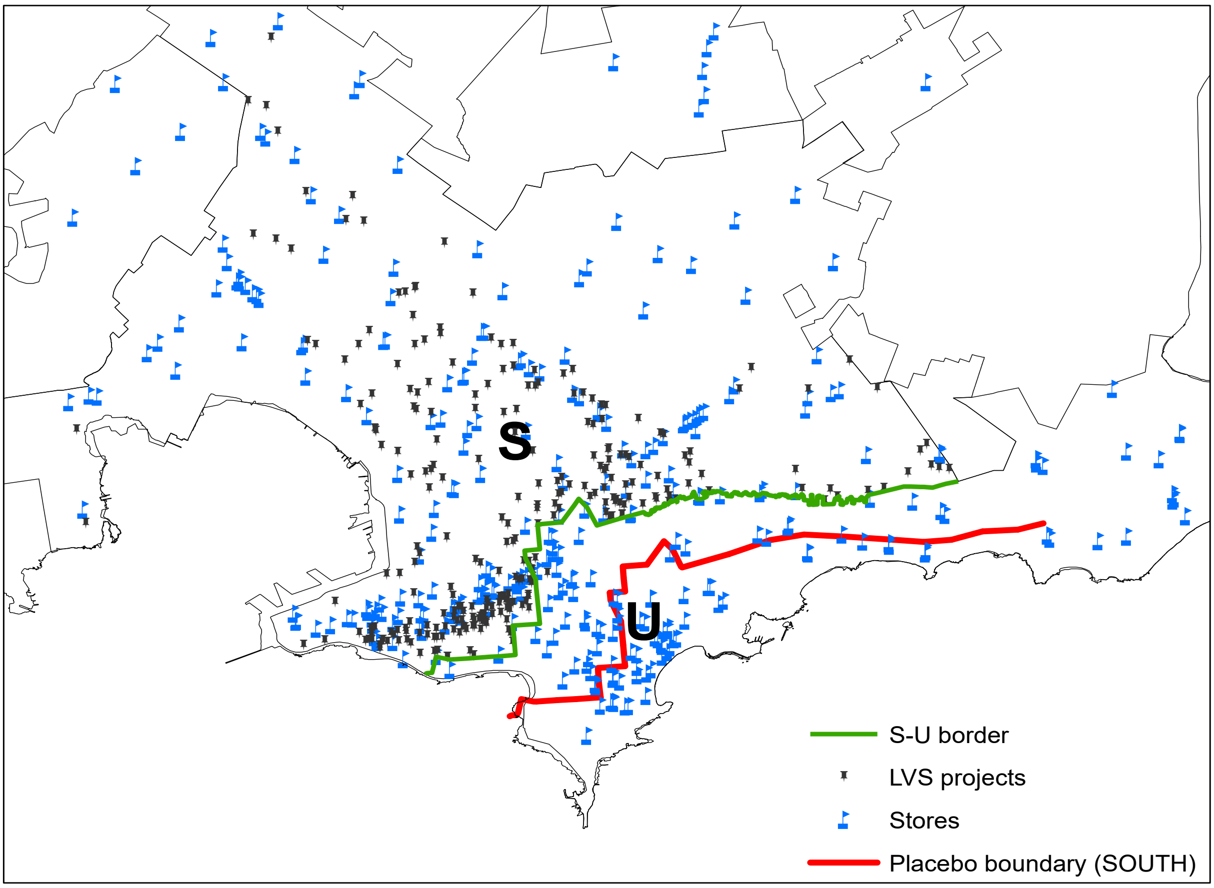
FIGURE A.8

PLACEBO EXERCISE NORTH



Notes: The placebo boundary resulted from shifting the southern border ($S - U$ border) to cross the centroid of the LVS area.

FIGURE A.9
PLACEBO EXERCISE SOUTH



Notes: The placebo boundary resulted from shifting the southern border ($S - U$ border) to the mid-point of the unsubsidized area.

TABLE A.8
PLACEBO - VARIETIES (REDUCED-FROM ESTIMATES)

	(1) Varieties Share (%)	(2) Varieties Share (%)
Post × Placebo	-3.780	4.133
	(7.496)	(4.554)
Placebo	South	North
Obs.	1249	781

Notes: Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

B. Theoretical Appendix

The Lagrangian associated to the consumer problem is given by

$$\mathcal{L} = q_0 + \alpha \sum_j q_i - \frac{1}{2}\gamma \sum_j (q_i)^2 - \frac{1}{2}\eta \left(\sum_j q_i\right)^2 + \lambda \left[y - q_0 - \sum_j p_j q_j\right]$$

From the FOCs with respect to q_0 we obtain $\lambda = 1$, while from the FOCs for variety j we obtain $\frac{\partial \mathcal{L}}{\partial q_i} = 0 = \alpha - \gamma q_i - \eta \sum_j q_i - \lambda p_i \Longrightarrow p_i = \alpha - \gamma q_i - \eta Q$.

B.1. ***Proof of Proposition 1***

In the final stage - when choosing quantities for a fixed N - the monopolist's problem becomes:

$$\max_{\{q_j\}_{j=1}^N} \sum_{j=1}^{N} q_j \left[\alpha - c - \frac{\gamma q_j}{L} - \eta \frac{\sum_{i=1}^N q_i}{L}\right]$$

Taking first order conditions for all varieties we obtain:

$$L(\alpha - c) - 2\gamma q_j - \eta q_j - \eta \sum_{i=1}^{N} q_i = 0$$

Given that, for an optimal choice of N, no q_j is equal to zero, these FOCs hold for all js. We can therefore solve for a generic j and obtain that in the symmetric equilibrium:

$$q^* = \frac{L(\alpha - c)}{2\gamma + \eta(1+N)} \qquad p^* = \frac{\alpha(\gamma + \eta) + c(\gamma + \eta N)}{2\gamma + \eta(1+N)}$$

Substituting these in the equation for profits in the varieties choice stage we obtain profits as a function of the number of varieties.

$$\pi(N) = \frac{L(\alpha - c)^2(\gamma + \eta)N}{(2\gamma + \eta(1+N))^2} - F_N N \tag{A.1}$$

To save on notation, we can re-write this expression as $\pi(N) = f(N) - F_N N$, where $f(N)$ is the first term in the right hand side of A.1. It is worth noting that the derivative of $f(N)$ is strictly decreasing in N, so the problem is concave. Therefore, it suffices to define the profit maximizing number of varieties N^* as the N that satisfies the condition $\pi(N) > \max\{\pi(N+1), \pi(N-1)\}$.

We now show that the number of varieties increases with market size L. Formally, this means that with L_1 and L_2 such that $L_2 > L_1$ – then $N^*(L_2) > N^*(L_1)$ where $N^*(.)$ is the optimal N for a given value of L. Define $\Delta(N) \equiv f(N) - f(N-1)$. Note that, because $f(.)$ is continuous and its derivative is decreasing in N, the function $\Delta(N)$ is also decreasing in N.

Given these conditions we can write the following system of inequalities:

$$L_2[\Delta(N^*(L_2))] - F_N > 0 \tag{A.2}$$

$$L_1[\Delta(N^*(L_1))] - F_N > 0 \tag{A.3}$$

$$L_1 << L_2 \tag{A.4}$$

Where the first and second conditions derive from the definition of $N^*(L)$ and the third is true by construction. Proceed by contradiction. Suppose that $N^*(L_1) = N^*(L_2)$. If this were the case, then – for low enough L_1 –either A.2 or A.3 need to be false, as the lower value of L_1 reduces the value of the positive component of A.3. Suppose instead that $N^*(L_1) > N^*(L_2)$. The fact that $\Delta(N^*(L_1))$ means that this would result again in a contradiction as the reduction from L_2 to L_1 is coupled with a reduction in $\Delta(N^*(L_1))$. Therefore, it has to be true that $N^*(L2) \geq N^*(L1)$ for $L_2 > L_1$.

It remains to show that this increase in varieties results in a reduction in prices. This is straightforward to see in the expression on p^* above, which is decreasing in N for the parameter restrictions outlined in the main text.

■

B.2. ***Proof of Proposition 2***

In the final stage, when choosing quantities, the first order conditions of firm m's problem can be written as:

$$L(\alpha - c) - \gamma q_j^m - \gamma \sum_{k=1}^{M} q_j^k - \eta \left(q_j^m + \sum_{k=1}^{M} \sum_{i=1}^{N} q_i^k \right) = 0$$

Define $Q_j \equiv \sum_{k=1}^{M} q_j^k$ and $Q \equiv \sum_{k=1}^{M} \sum_{i=1}^{N} q_i^k$. If we add the first-order conditions across firms first and then across varieties (js) we obtain:

$$M\left(L(\alpha - c) - \gamma Q_j - \eta Q\right) = (\gamma + \eta) Q_j$$

$$NM\left(L(\alpha - c) - \eta Q\right) = (\gamma + \eta + \gamma M) Q$$

Using these two expressions we can solve for Q, Q_j and q_j^m. Moreover, replacing the equilibrium value of q_j^m on demand we can obtain equilibrium prices. The resulting equilibrium expressions for quantities and prices are:

$$q^* = \frac{L(\alpha - c)}{\gamma + \eta + \gamma M + \eta NM} \qquad p^* = \frac{\alpha(\gamma + \eta) + c(\gamma M + \eta NM)}{\gamma + \eta + \gamma M + \eta NM}$$

Substituting these expressions in the firm's pay-off function we can obtain the expression for profits net of entry costs:

$$\Pi(M) = \frac{NL(\alpha - c)^2(\gamma + \eta)}{\gamma + \eta + \gamma M + \eta NM} - F - F_N N \tag{A.5}$$

The equilibrium number of firms is given by $M^* : \Pi(M^*) > 0, \Pi(M^* + 1) < 0$. Note that, an increase in L (keeping N fixed) can have two outcomes: either M^* stays the same or it increases. Re-writing $\Pi(M^*(L)) = Lg(M) - F - F_N N$ we know that:

$$L_2 g(M^*(L_2) + 1) < F + F_N N$$
$$L_1 g(M^*(L_1) + 1) < F + F_N N$$

Suppose $L_2 >> L_1$. In that case, we must have that $M^*(L_2) > M^*(L_1)$, otherwise (for sufficiently large gap between L_2 and L_1, either the first or the second inequality will not be satisfied. This proves that, for a fixed number of varieties, a large enough change in market scale L will lead to a larger number of firms in equilibrium. It is straightforward to see that this will result in a lower value of p^*, as long as $\alpha > c$.

■

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