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Local Retail Prices, Product Varieties and Neighborhood Change

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Abstract

We study how differences in retail prices within a city are affected by changes in local housing markets. Our empirical strategy is based on an exogenous shift in the spatial distribution of construction activity induced by a large-scale, place-based tax exemption in the city of Montevideo. We provide differences-in-differences and instrumental variable estimates showing that the price of retail goods decreases in areas within the city that experience more residential development. We use a multi-product model of imperfect competition to relate this change to an expansion in either product varieties or firm entry. We report evidence in support of the varieties channel, with new development causing an increase in varieties available locally. Our results have implications for urban planning policy and the broader discussion about winners and losers from neighborhood change.

JEL Classification: R23, R30

Keywords: Retail prices, Housing Supply, neighborhood change

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Local Retail Prices, Product Varieties and Neighborhood Change

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Abstract

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Keywords: Retail Prices, Housing Stock, Neighborhood Change.
JEL classification: R23, R30, R58

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1. Introduction

The availability of local retail options at grocery stores or supermarkets is not homogeneous within cities. Differences in the local consumer base across locations can shape the availability of these outlets as well as the prices and varieties of goods sold. Therefore, physical changes in neighborhoods that influence this consumer base can affect local retail opportunities and, in turn, the well-being of incumbent residents. This can operate through (at least) two channels. In the first place, changes in stock may increase residential density - 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 housing often attracts affluent residents (Brueckner, 2011) with high willingness for this type of housing. Through both channels, the local demand for goods and the demand for different varieties may increase with the development of new stock.[1] The way these changes map onto equilibrium prices and availability of goods will depend on the response of supply to the local demand shock.

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 to local households, and the entry of new stores in affected neighborhoods. In doing so, we study how incumbent households are be affected by changes in the physical infrastructure around them via mechanisms operating outside of the market for housing services.

When trying to understand how new stock affects local retail conditions, an important problem arises because residential development is shaped by local demand and is therefore endogenous to neighborhood amenities in the cross-section and over time. We overcome this problem by exploiting a major housing policy intervention that induced a large re-location in the development of new stock within the city of Montevideo, Uruguay.[2] In doing so, we build on previous work in González-Pampillón (2019) on the housing markets spillover effects of new development. The policy provides tax benefits to developers building housing in a pre-defined middle-income area of the city, effectively subsidizing development in those lo-

[1]The link between retail access and neighborhood change has been mainly studied in the urban planning literature on neighborhood change and gentrification. See for example, Meltzer and Capperis (2017), Mermet (2017), Zukin et al. (2009).

[2]Housing supply elasticities occasionally used in the literature to construct instruments for supply are usually calculated at the city level (see Saiz (2010) and Cosman et al. (2018)) cannot provide exogenous variation suitable for our analysis.

cations. Developers used the program intensely, with total investment through this program standing at a remarkable 1.5% of the GDP in the first five years of the policy. New units sold through the scheme were typically high-quality units in multi-family buildings marketed to mid-high/high-income households. We use this policy as an arguably exogenous shifter in the spatial distribution of residential construction in a strategy similar to a differences-in-differences design to induce exogenous variation in new stock around existing grocery stores and supermarkets. This strategy gives us an instrument to estimate the effect of new local stock on retail prices, varieties and entry.

We first test whether the introduction of new residential stock influenced the price of groceries available locally to consumers. Our estimates show that newly built housing resulted in a decrease in grocery prices at the local level. These effects are of moderate size and significant at conventional levels: Local grocery prices decrease by 2% in the areas affected by the policy relative to the comparison group. Our instrumental variable strategy allows us to re-scale this reduced-form coefficient to obtain an elasticity of prices to new stock between 3% and 4%. Thus, for fixed incomes, an increase in housing stock results in higher purchasing power for households in the vicinity of affected stores.

This result appears to be counter-intuitive, as local prices respond negatively to an increase in demand. 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 that framework, the increase in demand can lead to a reduction in prices 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: The varieties offered in supermarkets located in subsidized areas increased by an average of 12 percentage points due to the change in stock. The associated elasticity of varieties to newly built area amounted to 17%. In terms of entry, we find a transitory effect on the number of grocery stores available in affected areas but the effect largely dissipates a few years after the sale of subsidized units begins. 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: varieties increase substantially with a moderate reduction in the price of goods. These findings indicate that concerns about the detrimental equity effects of neighborhood change through a retail access channel may be misplaced.

Our analysis is carried out using a detailed good-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 detailed information from grocery stores all over the country. We also use official data from the National Housing Agency (*Agencia Nacional de Vivienda*) regarding subsidized new construction projects, which contains information about the exact geographical location of projects, approval date, total number of housing units produced, amount of taxes exempted and budget of each project, characteristics of built units, and project size (large, medium and small).

This paper is related to the growing literature on the effect of gentrification and neighborhood change on local outcomes. To our knowledge, none of these previous studies analyzes local retail prices. Various papers analyze residential mobility patterns in gentrifying tracts with a strong focus on displacement, finding mixed results. A group of studies that use more descriptive techniques find no (or mild) evidence on higher out-migration of original (and more vulnerable) residents while showing income gains among this neighborhoods (Vigdor, 2002; Freeman, 2005; McKinnish et al., 2010; Ellen and O'Regan, 2011a,b; Ding and Hwang, 2016). On the other hand, three recent studies (Aron-Dine and Bunten, 2019; Waights, 2018; Brummet and Reed, 2019) find that gentrification indeed lead to out-migration and displacement, while other study (Freeman et al., 2015) does not find evidence of higher mobility rates in gentrifying neighborhoods. Brummet and Reed (2019) also show that original residents who stay after the neighborhood gentrifies benefits from higher house values (whenever stayers are home-owners) and increased employment levels in the neighborhood. Vigdor (2010) empirically test if the willingness-to-pay to live in revitalized neighborhoods of existing residents exceeds the change in local rental prices (which captures changes in neighborhood quality). Autor et al. (2017) estimate the causal effect of gentrification induced by a rent deregulation policy on crime, finding a substantial reduction among crime rates. Closer to our work here, Asquith et al. (2021) study the effect of new residential stock on local prices and rents, finding a depression of local rents despite the new stock being occupied by relatively high-income residents.

Naturally, this paper is also related to the growing literature on urban consumption following the work in Glaeser et al. (2001). Some strands of this literature focused on endogenous consumption amenities (Diamond, 2016; Guerrieri et al., 2013; Almagro and Domınguez-Iino, 2019). Allcott et al. (2019) us a structural model of grocery demand to evaluate whether differences in the supply of groceries explains nutritional inequality in the United States. Perhaps closer to the question here, recent work in Couture et al. (2018) provides a model in which increasing inequality can interact with local consumption amenities and spatial sorting to make the poor worse off. Our paper looks 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.

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) estimates how changes in house prices affect local retail prices through housing booms and busts. 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 and which lead firms to increase mark-ups in response. We instead study how a process of physical change in certain neighborhoods affected local retail prices in stores in the vicinity.

2. Institutional setting

In August 2011, the Uruguayan government introduced launched 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).[3]. The LVS aims at promoting the construction sector and improving the stock of housing to be sold or rented by means of place-based tax benefits for new construction.[4] Developers and private investors building new stock in certain locations were exempted from paying corporate tax (25%) on profits made on the sale of sold units, while rents are partially exempted from personal income and corporate taxes for a period of 9 years.[5] 540 new construction projects were promoted from December 2011 until December 2019, involving almost 17K new units. The total amount invested was almost 1.4K MUSD, which is around 1.5% of the 2011-2019 average GDP in USD. The city of Montevideo concentrates 65% of the total projects (349 projects).[6] The average projects' schedule is 21 months.

Eligibility for the subsidy for new construction offered by the LVS policy was place-based. The relevant regions in Montevideo are observed in Figure 1. The tax benefit only apply 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) is the richest part of the city, with an average real

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

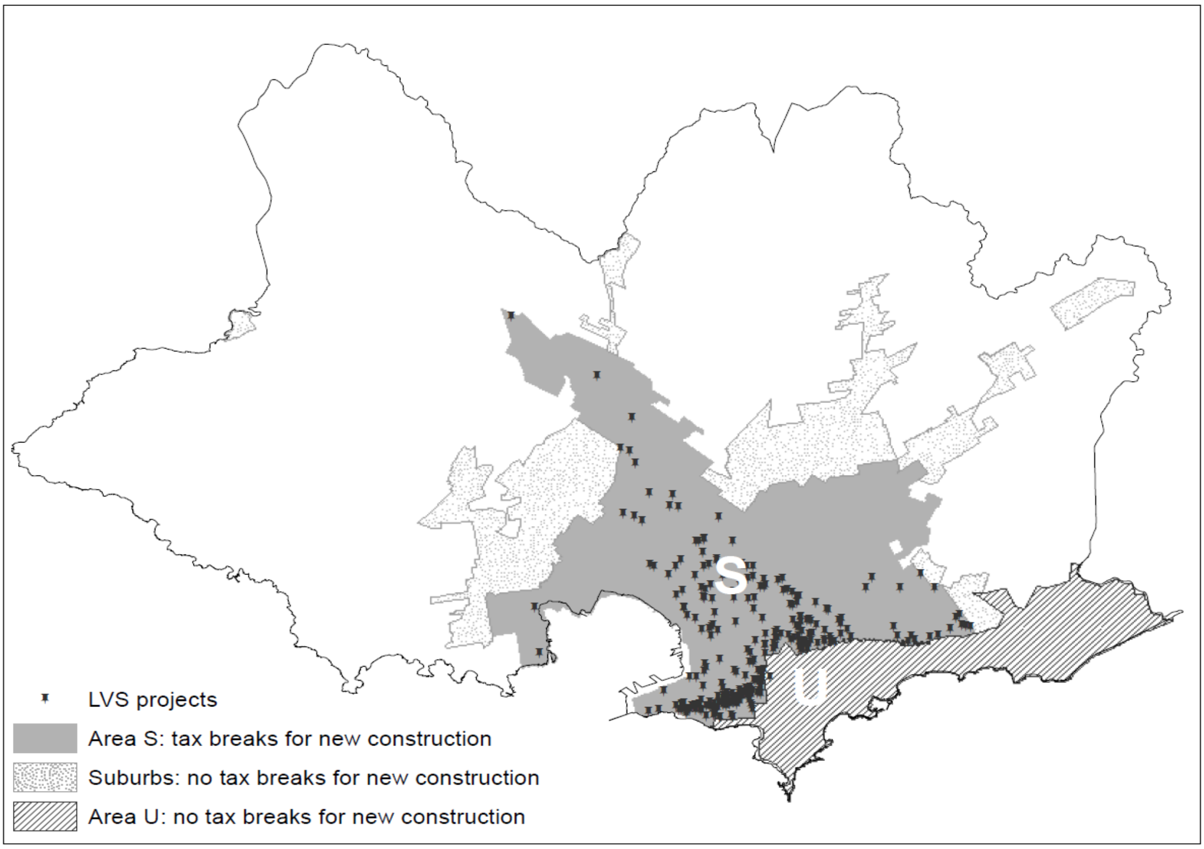
[4]The subsidy also applies for rehabilitation projects that consisted of upgrading and also increasing the total number of housing units. However, slightly above one fifth of the total comprises rehabilitation projects, having budgets substantially lower than the new construction ones.

[5]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.

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

per capita income that doubles and triples the one in area S and the suburbs (unlabeled), respectively. 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, showing a high concentration of projects on the southern border of area S. Three-quarters of the projects were performed within two kilometers of this border. This suggests that developers essentially chose locations near to the high price unsubsidized.

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.

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. There is no official document justifying the delimitation of the LVS area. Overall, it seems to follow some natural city divisions provided by its main avenues and streets. In addition, the policy seems to exclude the city's high-income areas and high-housing prices areas.

The LVS policy could be used to subsidize projects with a maximum of 100 new units by land lot. However, there were exceptions made for projects performed in large vacant lots or

[7] Appendix Figure A.2 in the appendix shows that the unsubsidized area U has a housing stock with better quality on average.

in parcels with disused factories or abandoned houses. Regarding unit size restrictions, it depended on the number of bedrooms (i.e. between $32m^2$ and $50m^2$ if one bedroom, increasing with each additional bedroom up to four).[8] The LVS units had to adhere to the guidelines laid down in the National Housing Plan and other ministerial regulations on quality. Figure A.1 (in the appendix) shows the distribution of quality for the LVS units and the existing stock. Around 95% of the LVS units were assessed as 'Excellent' by the Cadaster Agency, while the average non-LVS dwelling for the city is assessed as having regular quality.[9] The tax benefits only operated in urban areas with the exception of those with a high proportion of secondary dwellings.

3. Data

Our main dataset is based on a detailed good-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.[10] 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 in a tax law passed by the Uruguayan legislature in 2006, which changed the tax base and rates of the value-added tax (VAT). The Ministry of Economy and Finance was concerned about incomplete pass-through from tax reductions to consumer prices and hence decided to collect and publish the prices in different grocery stores and supermarkets across the country. The DGC issued Resolution Number 061/006, 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 objective of the DGC is to ensure that prices posted on the DGC website reflect the real 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.

Our 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

[8] The latest regulation increased the lower bound of one-bedroom LVS units to $35m^2$.

[9] The quality scale goes from 'Very poor' to 'Excellent'.

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

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 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 soft drink of the international brand Coca Cola is reported in its 1.5 liter variety by all stores. If this specific variety is not available at a store, then no price is reported. The data are then used 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.[11]

The three best-selling brands are reported for each market, disregarding the supermarket's brands.[12] Products were selected after a survey to some of the largest supermarket chains in the year 2006. In November 2011, the list of products was updated, including some markets and reviewing the top-selling brands for others. The price information for the discarded goods was deleted from the database, so we lose part of the information in some markets. The 154 products in the database represent more than 60 markets defined at the product category level (e.g., sunflower oil and corn oil and wheat flour 000 and wheat flour 0000 are different markets in our analysis). For some of them, the information does not identify the goods at the UPC level; in the meat and bread markets, products do not have brands. Of the 154 products, we identify 127 that could be exactly matched. 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. For the main analysis, we do not consider 54 products that entered the database in 2010-2011 as we do not have prices before the change of policy, but we test the sensitivity of our results to not dropping them.

For each supermarket, we have detailed information about the exact location given by its Universal Transverse Mercator (UTM) and whether it belongs to a chain. The database has information for up to 444 supermarkets - i.e., a non-balanced panel across all nineteen political states, comprising 54 cities. Montevideo, the capital city of Uruguay, is also the country's largest city[13], with nearly forty percent of the Uruguay population has 249 supermarkets in the sample.[14] See Borraz et al. (2014) for a completed description of the supermarket industry in Montevideo.

Our final database has 73 products corresponding to 33 markets/categories in 249 super-

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

[12]Exceptions are sugar, crackers, and cocoa, which has only two brands; and rice, which has up to six brands.

[13]More information is available at http://www.ine.gub.uy/uruguay-en-cifras.

[14]We do not consider supermarkets in the rural parts of Montevideo.

markets in Montevideo. We drop the year 2007 due to not having the first three months as it started in April.[15] We then calculate the mode monthly price (see Eichenbaum et al. (2011)) for each product. Our final database for the city of Montevideo is composed of 1,485,677 observations.

4. Residential Development and Retail Prices

4.1. Empirical Strategy

The primary aim of our empirical analysis is to estimate the effect of demand changes induced by new residential developments on local grocery prices. By local grocery prices we mean the price of groceries sold by local retailers (i.e. supermarkets) relative to some common benchmark. In our case, we will study how relative changes in residential supply across locations affects relative prices for sold goods.

One important identification problem when trying to detect this hypothesized causal link is that local demand for housing space in a given location – which will affect the location of new development – can itself be affected by local retail prices and the mix of local varieties available to consumers. In addition, local retail prices will be affected by the type of supermarkets available locally or the density of stores which, in turn, will be shaped by local urban planning decisions that also affect housing supply. In order to deal with these issues, we exploit the change in the spatial distribution of new residential development in Montevideo induced by the LVS policy to estimate our effect of interest.

We carry our analysis of price effects at the level of individual supermarket-product-time pairs. In the first place, we compare retail prices inside and outside the policy area and study its evolution over time before and after the changes in housing stock induced by this policy. We look specifically at areas within a two kilometer band of the LVS boundary because this is where we expect most of the change in the distribution of new development to take place. These areas are also more comparable than areas in the urban periphery (see Appendix Figure A.6). The reduced-form estimating equation in this case can be written as:

$$Log(P_{ist}) = \beta_{RF} Policy_s \times post_t + \alpha Policy_s + \delta_{it} + \epsilon_{ist} \tag{1}$$

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.

[15]Again, we test the sensitivity of our results to not dropping 2007.

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.

In the second place, we can use variation between sides of the LVS policy boundary over time as an instrument for (LVS and non-LVS) housing construction activity. To do so, we define a variable New Area$_{st}$ that measures the sum of the surface areas (in m^2) of new units within 1km of supermarket s. The variable is constructed using the accumulated stock of new units within six years (i.e., between $t-6$ and t) of t.[16] It measures the exposure of each supermarket s to new residential construction and, therefore, to a shift in local demand for its goods. We estimate the effect of New Area$_{st}$ on local retail prices by estimating the following via two-stage least squares:

$$Log(\text{New Area}_{st}) = \pi Policy_s \times post_t + \eta Policy_s + \omega_{it} + u_{ist} \quad (2)$$

$$Log(P_{ist}) = \beta_{IV} Log(\text{New Area}_{st}) + \delta_{it} + \alpha Policy_s + \epsilon_{ist} \quad (3)$$

where equation 2 is the first-stage and 3 is the second-stage a two-stage least squares (2SLS) estimation. 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.

The $Log(\text{New Areas}_{st})$ is instrumented by $Treat_s \times post_t$ used here as supply shifter. New housing construction in the treated area increased by 520K m^2 developed (almost 9K new units) between the pre- and post-policy period.[17] In Figure A.5 in the appendix, we use tract-level data and show that there is a discontinuity in new residential construction at the border, which we interpret as resulting from the incentives induced by the LVS policy. In addition to this descriptive evidence, we include estimates of first-stage equation 2 in Appendix Table A.3 and report the associated F-statistics in our tables.

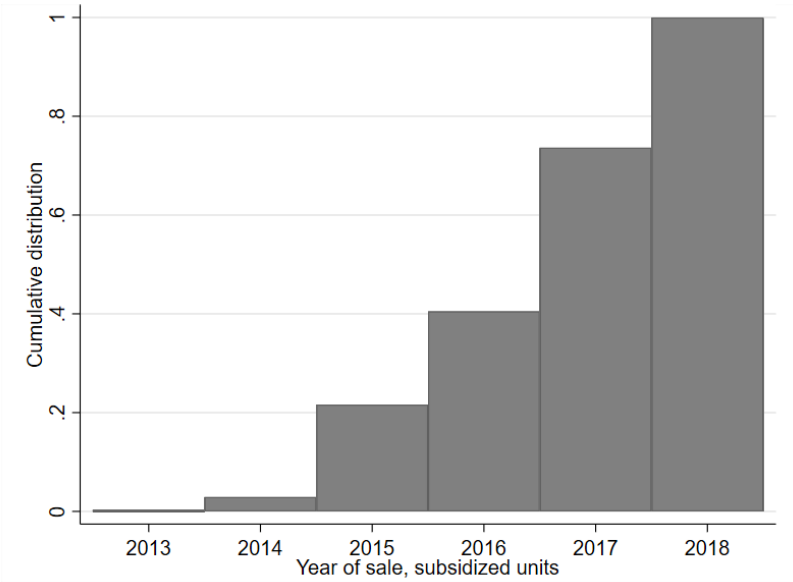
It is straightforward to see that both for the reduced-form and IV estimates the identifying variation is essentially the same: variation between regions across the boundary before and after the policy comes into effect. As a result, the identifying assumption is in both cases a typical parallel trend assumption as in a conventional difference-in-differences study. We

[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.

[17] These figures are based on our calculations using publicly available data from the Cadaster Agency. The pre-policy and post-policy period refer to 2004-2008 and 2014-2019, respectively. Construction activity remained at the pre-policy levels in the untreated area (i.e., 574K and 620K m^2 developed in the pre-and post-policy period), suggesting no crowding-out effect.

will show evidence in support of this assumption in the next section.

FIGURE 2
SALES OF LVS PROPERTIES



Note: Own calculations based on combining official data on LVS projects with data on housing transactions from the National Registry Office 2011-2018.

Variable $post_t$ in the equations above needs to be defined based on when new built induced by the LVS are sold. The accumulated sales of LVS projects started between 2011 and 2014 are displayed in Figure 2. We can see that very few sales had taken place before 2015, and the largest number of sales did not come until 2017. In order to capture the full effect of the new housing stock on local prices we will define $post_t$ as a dummy taking value one in 2019 and use price data for 2010 and 2019 in estimation. We discuss robustness of findings to changing the pre-policy year in section 6.

One additional element to take into account relates to the weighting of different products. We will report both unweighted, and CPI-weighted estimates in the following. Weighting is important because the effective price faced by households buying a bundle of goods depends on the relative size of each product in the household budget. We do not observe household consumption at the individual level. Therefore, 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 can use is use CPI weights of different product categories obtained from the National Statistical Office. Varieties of goods available may vary by supermarket, so we re-scale these weights in order to ensure that the joint weight of all products within a product category in a supermarket has a weight equal to the CPI weight ω_k^{CPI} for product category k. Hence, we compute $\omega_{is} = \omega_k^{CPI}/n_{kst}$ where n_{kst} is the number of products of category k offered in supermarket s in time t.

4.2. Results

Before turning to our estimates of the effect of new developments on prices we first show graphical evidence both to support our identification assumption and to illustrate our main finding. To do so, we estimate individual yearly coefficients obtained by replacing $\sum_{k=2008}^{2019} \rho_k Policy_s \times \mathbb{1}\{t = k\}$ instead of the interaction term in equation 1. Estimates of ρ_k capture the effect of the policy relative to the base year, 2010. These are reported graphically in Figure 3. The graph shows that the difference in grocery prices between stores in the treated and control bands around the LVS boundary is stable between 2008 and 2012. The p-value of a joint test for equality coefficients ρ_{2008} through ρ_{2012} is 0.984. This is reassuring as it suggests that the parallel trend assumption required for identification is satisfied in our data.

We observe coefficients continue to be statistically insignificant in subsequent years up to 2016. This is consistent with the fact that only a relatively small fraction of new LVS units had been effectively sold before 2017 – as shown in Figure 2 – so that neither the local population density nor the composition of this population had been effected much by the policy yet. In 2017 that we find a clear break in the trend towards the larger effects around -3% that we observe in 2018 and 2019.[18]

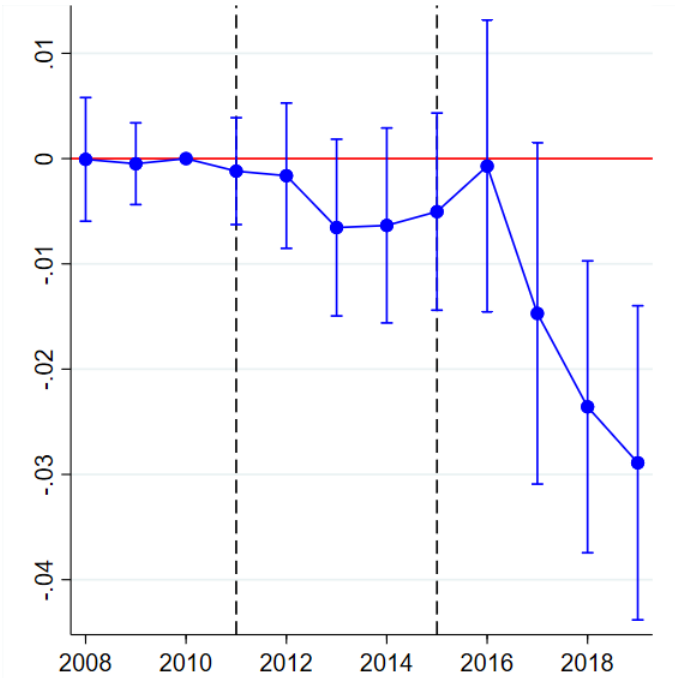
Quantitative estimates of the effect of new stock on prices are reported in Table 1. Columns 1 and 2 provide reduced-form estimates – see equation 1 – and columns 3 to 6 provide IV estimates – see equation 3. Estimates reported in columns 2 and 4 are obtained using CPI product-category weights. We find *negative* and significant effects of new housing developments on prices across the board. Columns 1 and 2 indicate that grocery stores located in the subsidized side of the LVS boundary reduced prices by roughly 2% relative to those across the boundary.

Instrumental variable estimates of the effect of new residential supply on prices are reported in columns 3 and 4 of Table 1. The estimated elasticity of retail prices with respect new housing area ranges from -.031 to -.038 (with and without CPI weights, respectively).[19]

[18]In Appendix Figure A.7, we plot the event study graphs that result from using CPI product-category weights as described above, including the year 2007, including products that entered in the database in 2010/2011, and using product-brand-time effects to account for aggregate differences in product prices over time. The figure shows that the qualitative findings from the event study analysis for prices are not sensitive to these modeling choices.

[19]Our first-stage estimates reported in panel A of Appendix Table A.3 indicates 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 both

Figure 3
Event-Study Graph: Prices



Note: Reduced-form event-study type coefficients. Round markers indicate coefficients obtained from replacing $\sum_{k=2008}^{2019} \rho_k Policy_s \mathbb{1}\{t = k\}$ instead of the interaction in equation 1. Effects are relative to 2010 the omitted year. Vertical segments correspond to 95% confidence bands. Dashed lines corresponds to years 2011 and 2015.

Taken together, these results confirm the findings illustrated in Figure 3. The new developments resulted in a moderate reduction of grocery prices available to households living in these areas. 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 that household with relatively lower incomes spend more of 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.

5. Mechanisms: Varieties and Entry

From the point of view for the local market for groceries, we interpret the change in housing stock induced by the LVS policy as increasing local demand in the treated areas. Under this interpretation, our finding that grocery stores near the new developments reduced good

the specifications with and without weights.

Table 1

Reduced-form and IV Estimates - Grocery Price Effects of New Developments

	Reduced-Form		IV	
	(1)	(2)	(3)	(4)
Policy × Post	-0.024***	-0.020**		
	(0.008)	(0.008)		
Log(New Area)			-0.038**	-0.031**
			(0.016)	(0.015)
CPI Weights	N	Y	N	Y
1st F-stat			21	21
Obs.	132192	132192	132192	132192

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

prices appears counter-intuitive. A conventional supply and demand framework would make the opposite prediction in the face of an increase in demand. Yet, this conventional framework may not be appropriate in this context. This is the case because supply itself may respond to the changes in demand. This can operate via at least two channels: i) entry can lead to non-linear or lumpy changes in the number of suppliers can increase competition and reduce prices, or ii) the introduction of new varieties may change the pricing incentives of supermarkets and prompt a reduction in the price of existing varieties. We turn to explore these mechanisms below. We first present a model featuring these mechanisms and then analyze these channels empirically by following a strategy similar to that used for prices.

5.1. 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. When the number of consumers increases, the threat of entry for established stores implies reducing markups to deter new entry. To achieve this result, established firms increase the number of varieties offer – which allows price discrimination of consumers – and at the same time has a downward effect on prices. The decrease in prices reduces the expected profit of entrants and, as a result, discourages entry.

To rationalize how an increase in local demand can result in lower prices, we propose a model based on Mayer et al. (2014) that features both the entry and product variety channels. In the model, 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:

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

where q_0 and q_i represents the individual consumption of the *numeraire* good and each variety i, respectively. The demand parameters α, γ, and η are all positive. Maximizing utility we obtain the individual inverse demand for each variety:

$$p_j = \alpha - \gamma q_j^c - \eta Q. \tag{4}$$

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 $q_j^c = \dfrac{\sum_{k=1}^{M} q_j^k}{L}$, 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:

$$p_j = \alpha - \gamma \frac{\sum_{k=1}^{M} q_j^k}{L} - \eta \frac{\sum_{k=1}^{M} \sum_{j=1}^{N} q_j^k}{L} \tag{5}$$

Firms face entry costs F, fixed costs of offering each variety F_N and fixed marginal costs per unit c, with $c < \alpha$.[20] When considering the multi-firm equilibrium, we consider firms enter sequentially. Entrants then simultaneously choose the varieties to be produced, and then they simultaneously choose quantities. Firm profits are therefore given by $\pi^m = \sum_{j=1}^{N_j} \left[q_j^m \left(p_j^m - c \right) - F - F_N N \right]$. 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:

$$\alpha - c - \frac{\gamma q_j^m}{L} - \frac{\gamma \sum_{k=1}^{M} q_j^k}{L} - \eta \left(\frac{q_j^m + \sum_{k=1}^{M} \sum_{i=1}^{N} q_i^k}{L} \right) = 0 \tag{6}$$

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

[20] 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.

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.

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 price of sold goods. Thus, we show that an expansion in the market for a retailer can lower prices via an expansion in varieties.

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 p^*, so that an increase in demand can lead to lower prices via its effects on entry, even if the number of varieties is fixed.

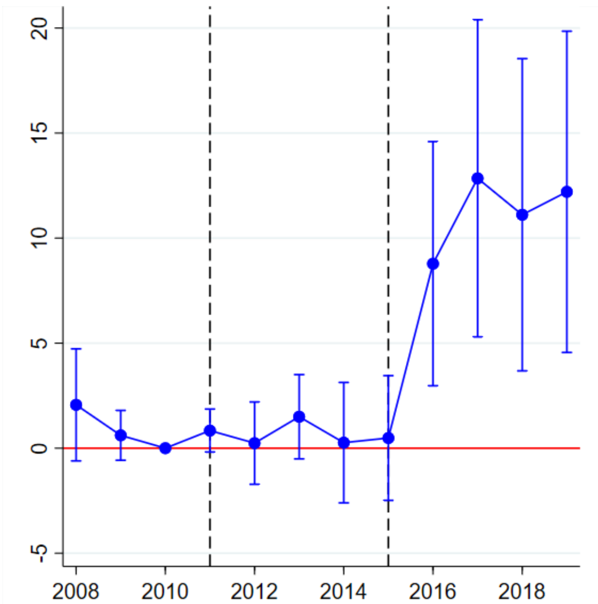
We have shown that both changes in varieties available or entry can provide scope for a reduction in prices resulting from a change in demand. Which of these mechanisms was behind our baseline results for the effect of new building activity in Montevideo? We turn to this question in the following sections.

5.2. Empirical Evidence: Change in Varieties

Informed by the results in the previous section, we now turn to test whether the change in local demand resulted in an increase in the varieties available to consumers locally. For this purpose we will exploit the same empirical strategy used in section 4, relying on exogenous variation induced by the shift in construction activity within the city. We measure varieties at the supermarket level, by calculating the fraction 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 3, using our measure of varieties as an outcome in a grocery store panel with interacted year effects. Coefficients for these interaction terms are illustrated in Figure 4, 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 2013. We cannot reject the null that the coefficients for 2008 through 2013 are equal (p-value 31.6%). A substantial change is observed starting in 2016. Note that this coincides with the period in which we observe the break for prices. The coefficients for 2016 through 2019 are positive and much 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. While confidence intervals for individual coefficients often cross the zero line, a joint significance test rejects the null (p-value: 1.23%).

FIGURE 4
EVENT-STUDY GRAPH: VARIETIES



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.

To obtain the DID estimates of the effect of the change in housing stock on available varieties, we estimate the modified version of equations 1, 2 and 3 using a store panel for the years 2008 and 2019.[21] DID estimates of the effect of new developments on varieties are reported in Table 2. Column 1 reports reduced-form estimates indicating that grocery stores in the side of the boundary that received the tax exemption experienced a relative increase in varieties of 9.1 percentage points. Column 2 and 3 report IV estimates using the number of new units and the sum of the squared meters of the new units, respectively. Results indicate that a one percent increase in these measures of proximity to new developments increases varieties by 17.6 and 14.5 percentage points, respectively.

We interpret these findings in light of the model presented in section 5.1. 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 2
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. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

5.3. Empirical Evidence: Entry

Changes in local housing stock may prompt an increase in local demand and lead to the entry of new grocery stores in affected neighborhoods. This could reduce local retail prices for residents, as shown in our theoretical framework. To investigate whether this mechanism could explain our finding above, we estimate the effect of the change in housing stock on access to grocery stores at the local level.

For this purpose we first compute variable:

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

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

.

$$\text{Grocer Access}_{ct} = \sum_{s=1}^{S} \frac{D_{st}}{d_{cs}} \tag{7}$$

Grocer Access$_{ct}$ is an inverse distance weighted average of access to grocery stores computed for each census tract c in every year t. S is the total number of stores in our sample (249). 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. We see Grocer Access$_{ct}$ as measuring local access to grocery stores. In addition to using inverse distance weighting as above, we also consider an alternative definition in which we only count grocery stores within 1km of each census tract.[22] Using these variable definitions in a census area panel covering the period 2008-2019 we estimate our reduced form equation:

$$\text{Grocer Access}_{ct} = \alpha_c + \delta_t + \beta_{RF} Policy_c \times post_t + \varepsilon_{ct} \tag{8}$$

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 obtained when using different proxies for access to grocery stores are reported in Table 3. We report both reduced-form and IV estimates. The outcome variables in each column are: the log of the number of stores within 1km in column 1, 3, and 5, the log of expression in equation 7 for inverse distance weighted access to grocery stores in column 2, 4, and 6.

TABLE 3
REDUCED-FORM ESTIMATES - GROCERY STORE ENTRY

	Reduced-Form		IV	
	(1) <1km	(2) 1/d	(3) <1km	(4) 1/d
Policy $\times$ Post	0.027 (0.030)	-0.008 (0.009)		
Log(New Area)			0.031 (0.034)	-0.010 (0.010)
First-stage F-stat			240	240
Obs.	852	854	852	854

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

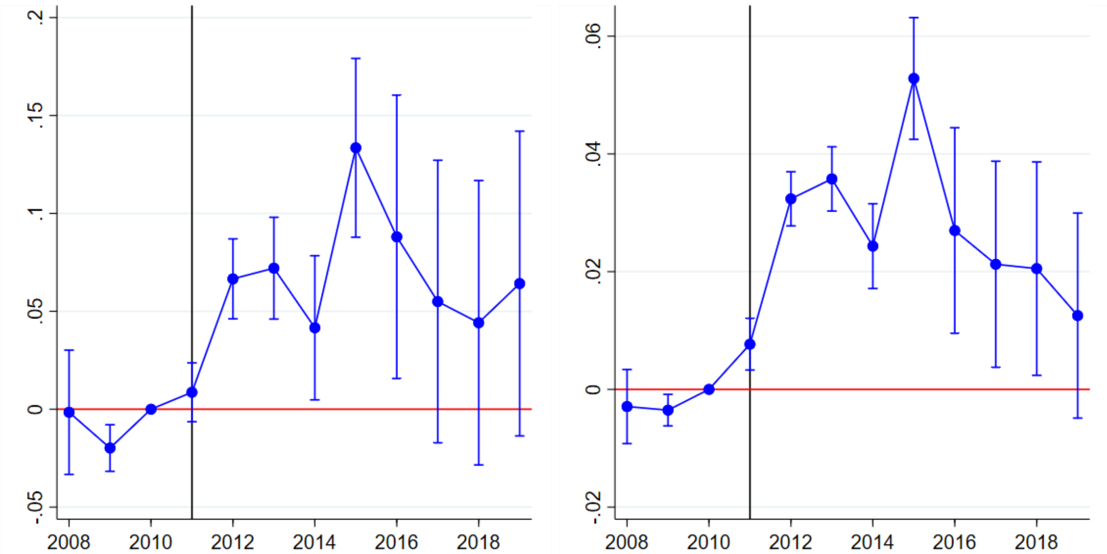
Results in Table 3 would lead us to conclude that the changes in housing stock induced by the LVS did not lead to an increase in entry. Yet this conclusion is slightly misleading.

[22] In that case we would have Grocer Access$_{ct} = \sum_{j=1}^{S} D_{st}\mathbb{1}\{d_{cs} < 1km\}$.

Figure 5 plots an event study type graph of the effect of the reduced form effect of $Policy_c \times \delta_t$ on the log of the number of stores within 1km of each census area (top panel) and the log of accessibility measure using inverse distance decay weights. We can observe that the introduction of the LVS policy did lead to entry initially, with access to grocery stores increasing after 2011 in the treated relative to the untreated area. 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 3.

We interpret this finding as suggesting that the (anticipated) change in stock led to a reshuffling of the types of grocery stores operating in the area. This dynamic aspect of the change in local stores is not incorporated in our static theoretical framework. It indicates that the change in varieties does not come (exclusively) from a change in varieties offered by individual stores but is rather a response in the type of stores available to consumers. Hence, while the number of stores available to households in these areas only show a timid increase in the long-run, entry may have provided the adjustment margin for the change in varieties and the decline in prices.

FIGURE 5
EVENT-STUDY GRAPH: ENTRY



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

Note: Round markers indicate estimated coefficients from a regression of grocery shop access 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.

6. 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 theway we measure new building activity next to stores, ii) using an alternative baseline year, iii) fixing the number of varieties, iv) fixing the number of stores, and v) estimating price effects separately for leader and non-leader 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 and stores 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 1 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 4, 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 1) 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 consider new 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 8.9% to 17.6% compared to our baseline estimate of 14.5%, and being statistically significant at the 5% level in most the cases. The picture is similar in case of entry. Except for one point estimate that is statistically significant at the 10% level, the resulting estimates are not significant, indicating that the change in the number of stores between 2010 and 2019 was not concentrated in areas where new residential development was taking place.

In our main specification for prices, we allow both the stores and the available varieties to vary over time. Yet we can restrict our sample to a fixed number of initial varieties and stores to explore whether these decision has any influence on our qualitative findings. We do so in Panel A of Appendix Table A.6. We first fix varieties by looking at the price effect on products available in the pre-intervention bundle in 2010. Coefficients obtained in this way allow us to know whether the price effects reported in section 4 are due to a compositional effect as stores started offering products which were previously unavailable. Results are reported in columns 1 and 2 of Panel A of Appendix Table A.6. Interestingly, we still observe a reduction in retail prices with point estimates being slightly larger than those obtained when the varieties by store can vary over time.

We can alternatively fix the stores while allowing varieties by store to vary over time. To do so we keep data fir stores that were present in both years 2010 and 2019. We report results under this alternative sample restriction in columns 3 and 4 of Panel A of Appendix Table A.6. The idea is to test if the new stores are the ones driving prices down. Again, we observe the negative price effects reported above continue to be present for this sub-sample of stores.

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 Placebo´ area as the placebo policy region to test whether differences between these regions emerge in our outcomes of interest (see Figure A.10 for a graphical description). This first exercise is labelled 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 5, while results for varieties are presented in column 1 of Appendix Table A.7.

The second exercise – labelled as placebo *North* – is constructed by shifting the policy border northwards up to the centroid of the LVS subsidised area S (see Figure A.9 for a graphical description). In this case, we consider stores within two kilometers of the constructed border and 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 5, and for varieties are reported in column 2 of Appendix Table A.7. All placebos yield statistically insignificant effect and point estimates

Table 4
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: Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

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

that are substantially lower than those reported in our main analysis.

7. Conclusions

neighborhoods are shaped by the physical infrastructure in them, with a key role played by housing in particular. As a consequence, large scale changes in stock induce a process of neighborhood change. The results in this paper show that, when change is induced by large scale residential development activity, local retail conditions in terms of the price and varieties of groceries available at the local level change. We find evidence of a moderate *reduction* in prices as a response to this change in demand, as well as a substantial increase in available varieties for local residents. Using our model, we show that these two facts can arise in the context of multi-product firms competing in quantities. We show that, entry of new stores plays a role in promoting this change, yet we do not find robust evidence of a sustained increase in the number of stores available at the local level.

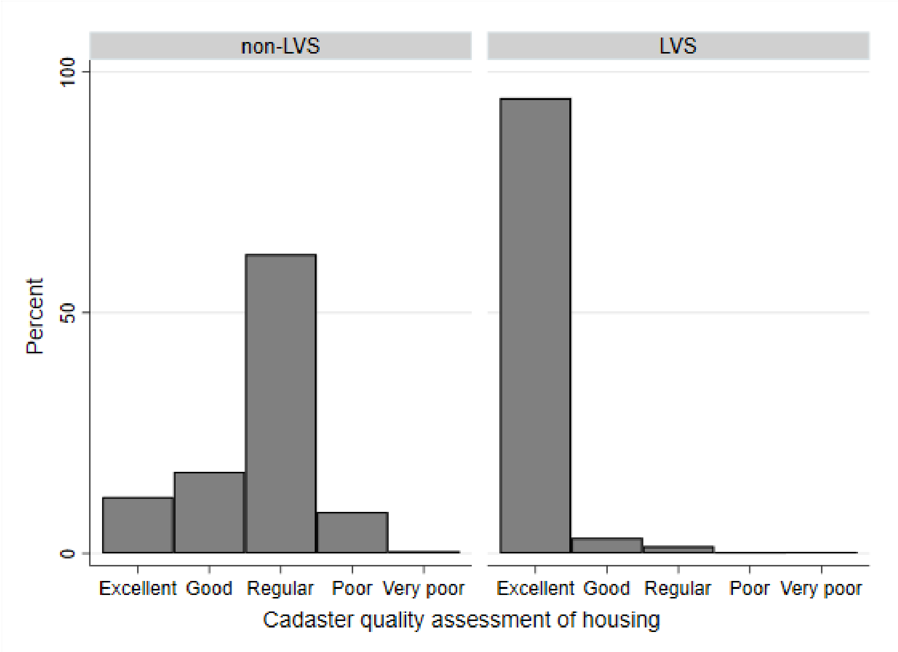
The combination of a reduction in prices and an increase in varieties for a largely fixed number of stores corresponds to a net improvement in the conditions for grocery consumers at the local level. Therefore, our results emphasize advantages of new development and neighborhood change for incumbent residents that had 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 and varied grocery products.

Appendices

A. Additional Figures and Tables

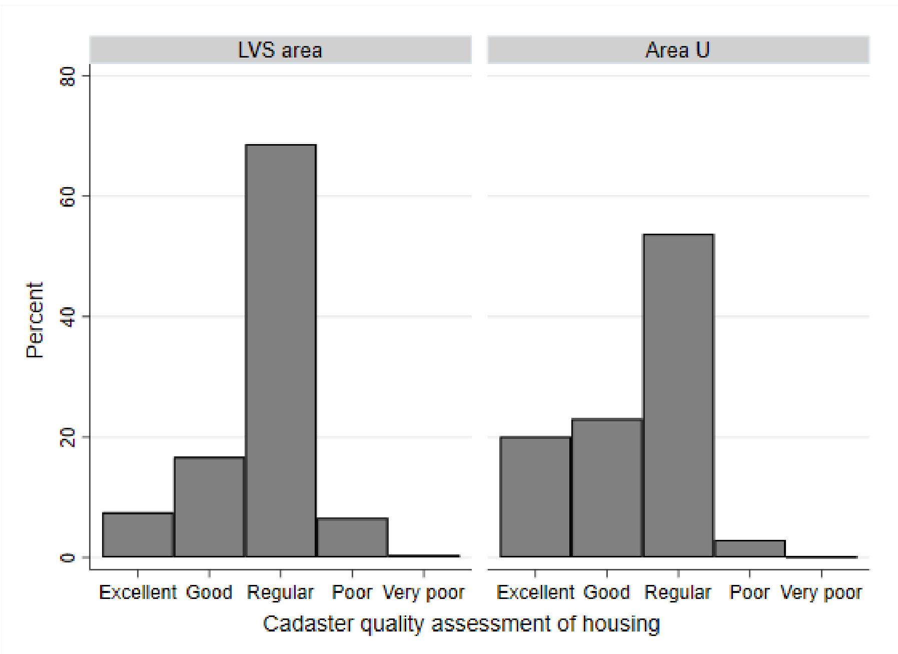
A.1. Quality of LVS units

FIGURE A.1
QUALITY OF LVS UNITS



Source: Own calculations based on data from the Cadaster Agency. *Notes:* The quality scale goes from 'Very poor' to 'Excellent'.

FIGURE A.2
QUALITY OF HOUSING WITHIN TWO KM OF BORDER $S - U$

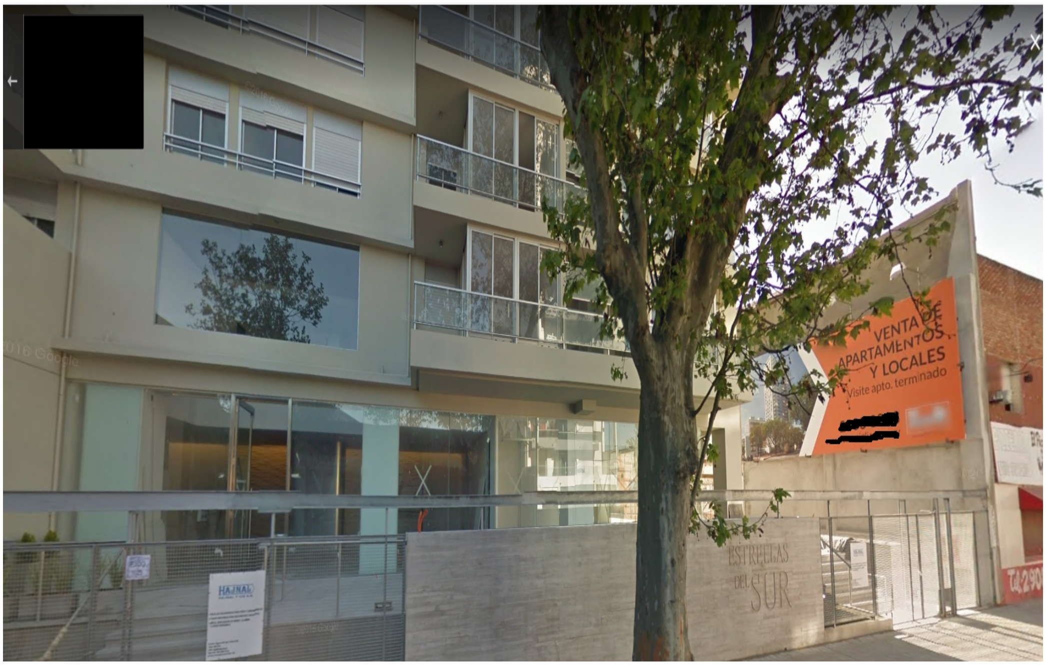


Source: Own calculations based on data from the Cadaster Agency. *Notes:* The quality scale goes from 'Very poor' to 'Excellent'.

Figure A.3
Example of a LVS project

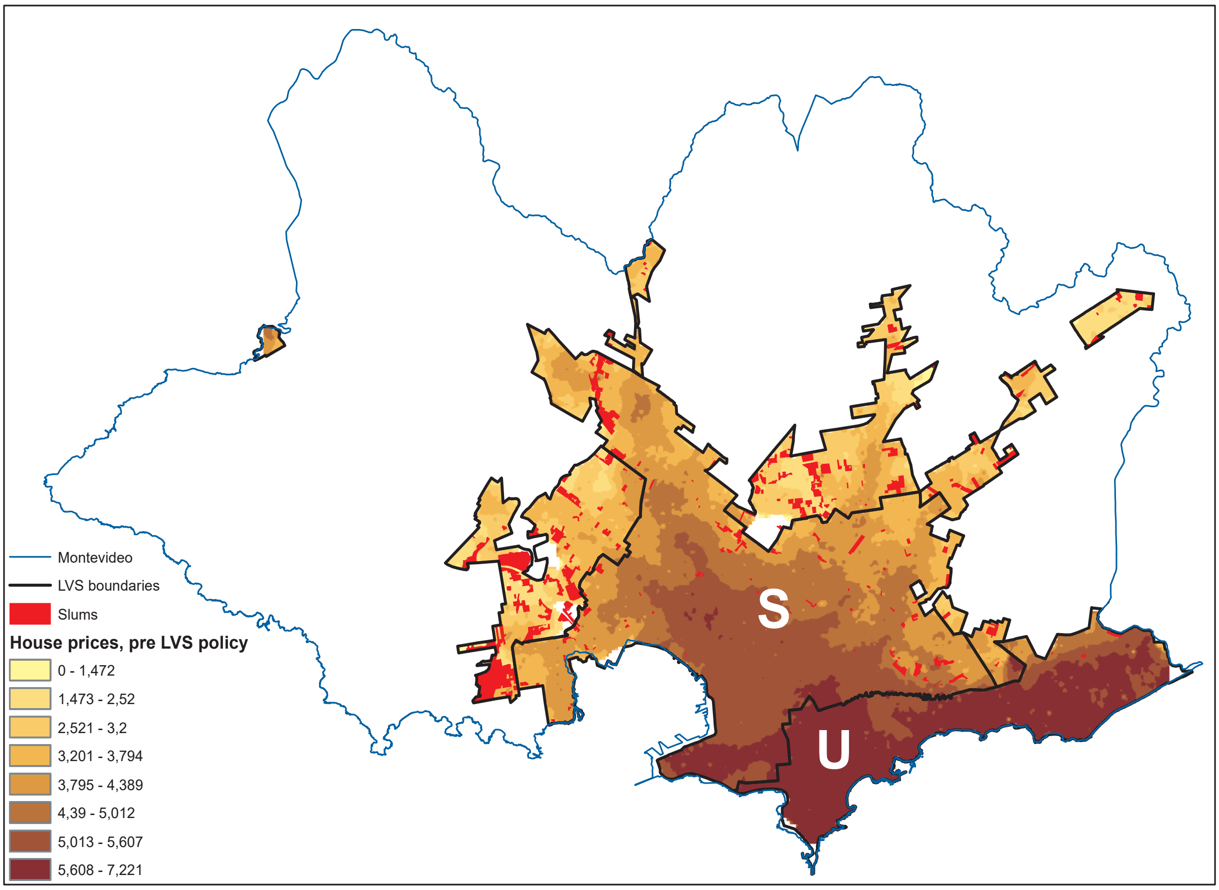
(a) Before

(b) After



A.2. House prices pre LVS policy

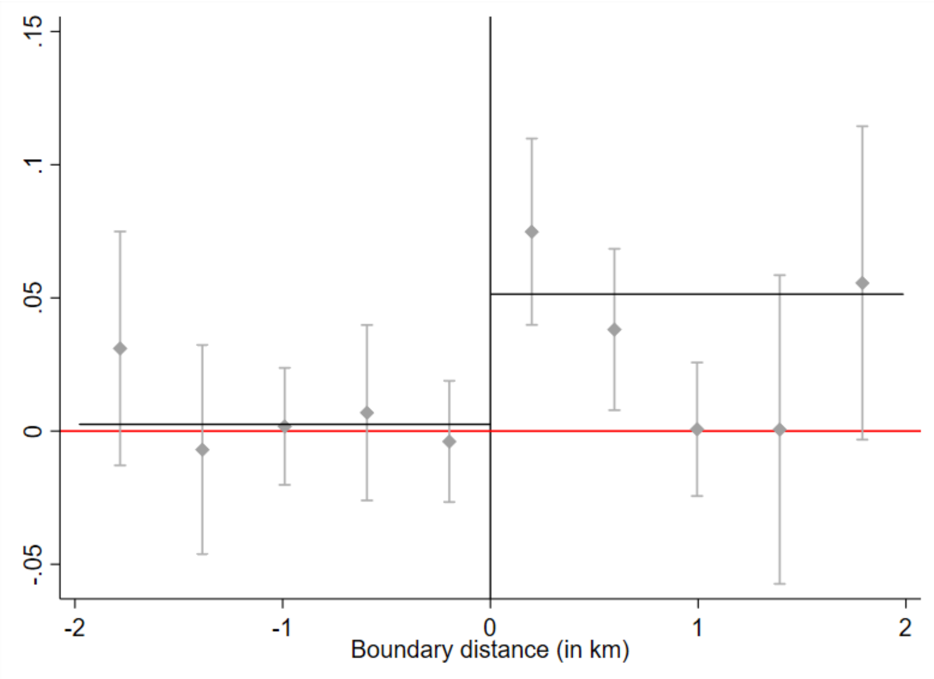
FIGURE A.4
MAP OF HOUSE PRICES (IN M^2, PRE LVS POLICY)



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.

Figure A.5

Changes in m^2 of new residential buildings



Note: Evenly spaced bins measure the average change in m^2 of new residential developments before-and-after the introduction of the policy (relative to the existing stock) at different distances from the border, using the method developed in Calonico et al. (2015, 2017). Negative (positive) distance values denotes locations in the treated (untreated) area. Vertical segments are 95% confidence sets. Fitted lines are estimated using a zero-order polynomial.

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

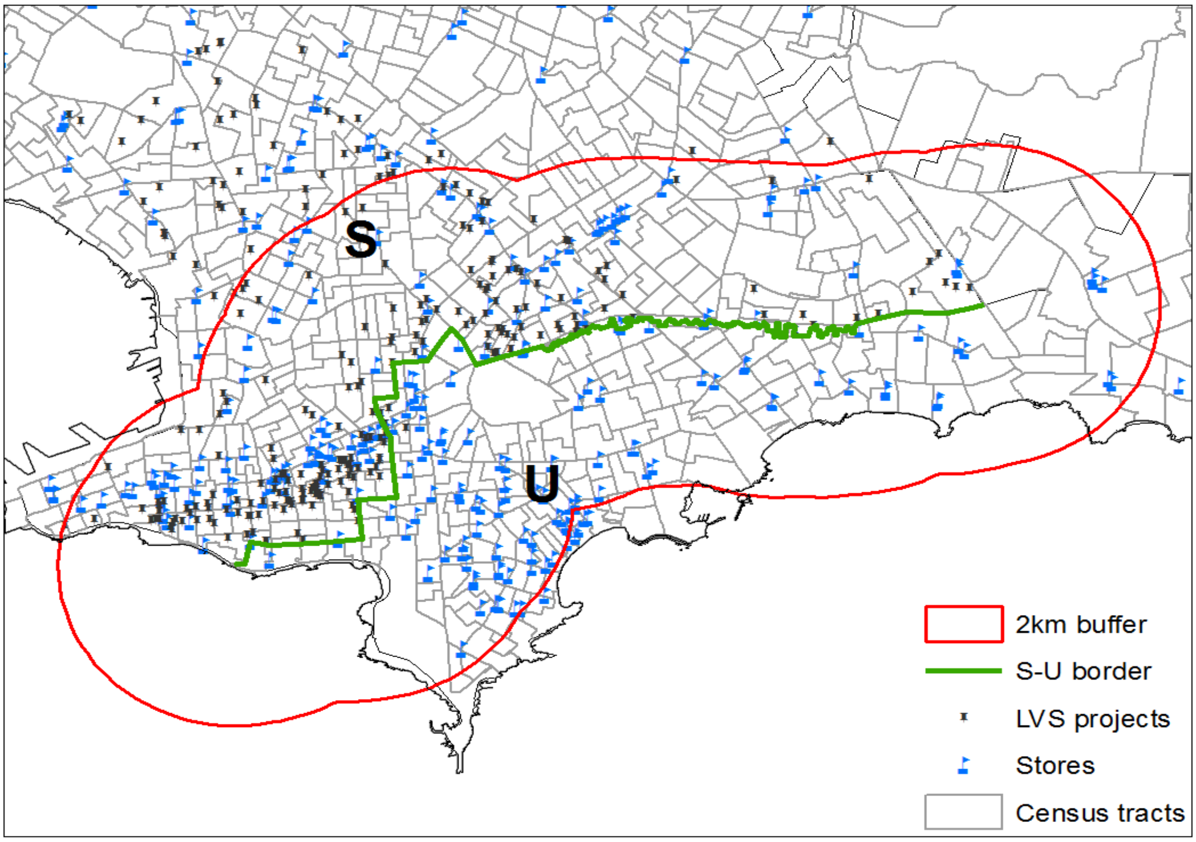
Table A.2
List of products (continued)

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.

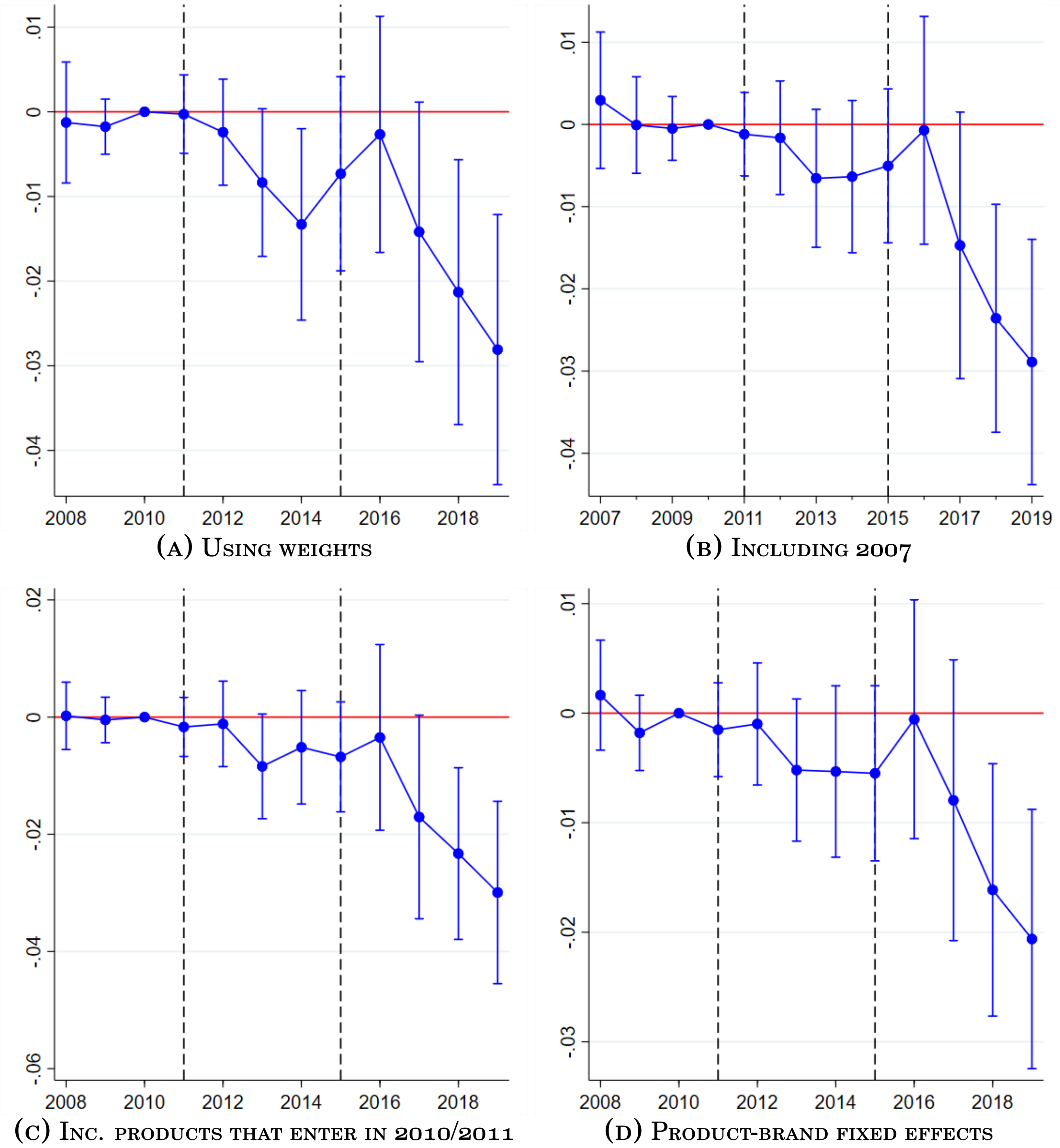
Figure A.6

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.7
EVENT-STUDY GRAPH: PRICES



(A) USING WEIGHTS

(B) INCLUDING 2007

(C) INC. PRODUCTS THAT ENTER IN 2010/2011

(D) PRODUCT-BRAND FIXED EFFECTS

Note: Round markers indicate coefficients obtained from replacing $\sum_{k=k_0}^{2019} \rho_k Policy_s \mathbb{1}\{t = k\}$ instead of the interaction in equation 1. a) $k_0 = 2008$; b) $k_0 = 2007$. Effects are relative to 2010 the omitted year. Vertical segments correspond to 95% confidence bands.

TABLE A.3

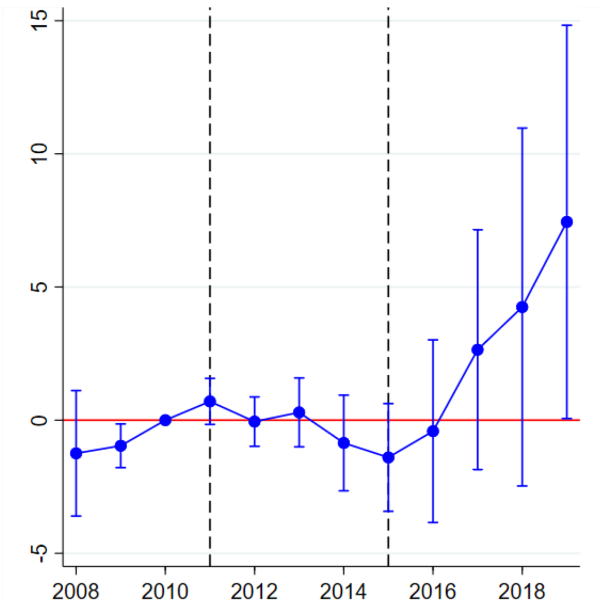
FIRST-STAGE - NEW DEVELOPMENTS EFFECTS OF THE LVS POLICY

	(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.8

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: 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)	0.035	-0.011		
	(0.040)	(0.011)		
First-stage F-stat	227	227		
Obs.	852	854		
	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)	0.029	-0.009	0.034	-0.011
	(0.033)	(0.009)	(0.039)	(0.011)
First-stage F-stat	220	220	221	221
Obs.	852	854	852	854
	New housing within 1.5km		New housing within .5km	
	<1km	1/d	<1km	1/d
C. Area Around Retail Store				
Log(New Area)	0.036	-0.011	0.026	-0.010
	(0.041)	(0.011)	(0.035)	(0.010)
First-stage F-stat	318	318	100	100
Obs.	852	854	848	850
	Baseline Year: 2008		Baseline Year: 2009	
	<1km	1/d	<1km	1/d
D. Alternative Baseline Year				
Log(New Area)	0.047	-0.004	0.063*	-0.005
	(0.038)	(0.010)	(0.035)	(0.010)
First-stage F-stat	178	178	228	228
Obs.	852	854	852	854

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

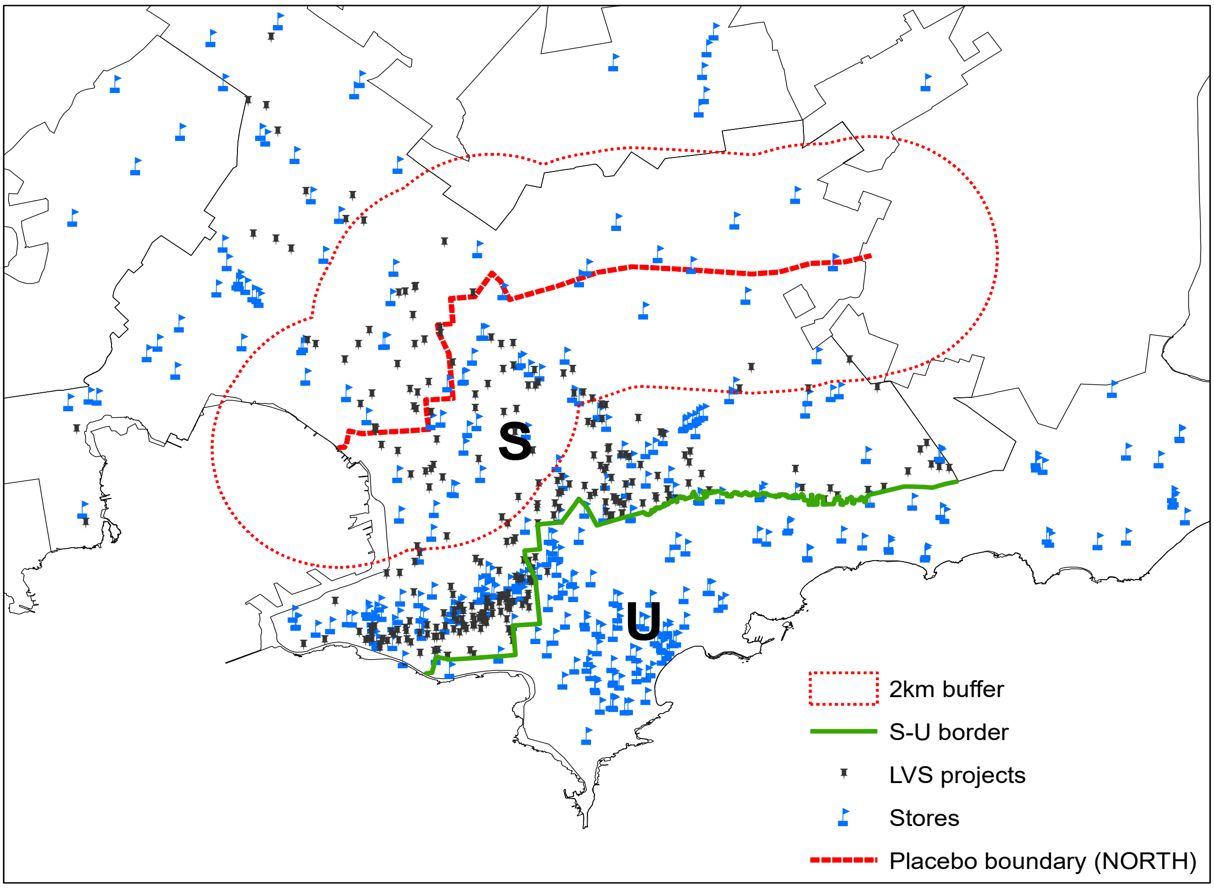
Table A.6
Additional Robustness Checks - Price Effects

	Fixing varieties		Fixing stores	
	(1)	(2)	(3)	(4)
A. Fixing varieties/stores				
Log(New Area)	-0.040***	-0.036***	-0.026**	-0.021**
	(0.013)	(0.012)	(0.012)	(0.010)
CPI Weights	N	Y	N	Y
First-stage F-stat	24	26	28	32
Obs.	115437	115437	107374	107374
	Leader brand		Non-leader brand	
B. Type of 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: Standard errors are clustered at the store level. *, **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Figure A.9

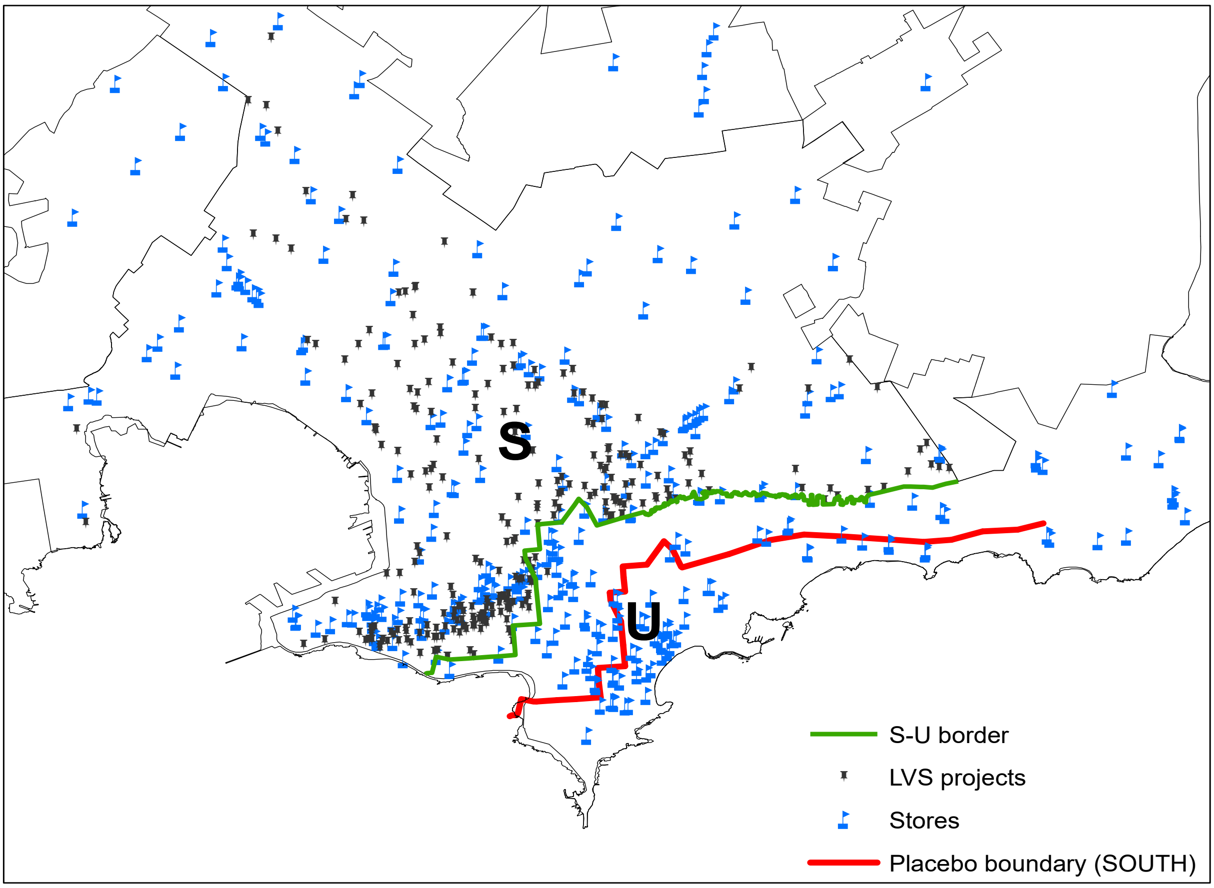
Placebo exercise NORTH



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

FIGURE A.10

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.7

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:

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

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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