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## The economics of the public option: Evidence from local pharmaceutical markets

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# The Economics of the Public Option: Evidence from Local Pharmaceutical Markets

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# The Economics of the Public Option: Evidence from Local Pharmaceutical Markets

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## Abstract

We study the effects of competition by state-owned firms, leveraging the decentralized entry of public pharmacies to local markets in Chile. Public pharmacies sell the same drugs at a third of private pharmacy prices, because of stronger upstream bargaining and market power in the private sector, but are of lower quality. Public pharmacies induced market segmentation and price increases in the private sector, which benefited the switchers to the public option but harmed the stayers. The countrywide entry of public pharmacies would reduce yearly consumer drug expenditure by 1.6 percent. (JEL D72, H4, I16, L3)

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State-owned firms compete with the private sector in education, healthcare, insurance, and basic services, among others. Supporters of the public option argue that it helps discipline markets that fail to provide enough incentives for private competition, because of either information asymmetries, market power, collusive behavior, or other market failures (Atkinson and Stiglitz, 1980). In contrast, critics argue that state-owned firms might be inefficient, provide low quality, or be captured by political interests (Shleifer and Vishny, 1994; Shleifer, 1998). Estimating the equilibrium effects of the public option has been difficult due to the lack of exogenous variation in the extent of public competition and the scarcity of contexts that allow evaluation of its distributional and political consequences.

In this paper, we study the decentralized and large-scale entry of public retail pharmacies in Chile, where pharmacies managed by local governments entered 146 of the 344 counties between 2015 and 2018. Public pharmacies emerged as nonprofit competition to a fully deregulated and highly concentrated private retail market characterized by high prices.<sup>1</sup> Public pharmacies sell drugs at prices that are 34 percent of those charged by their private counterparts. These low prices are possible because private pharmacies hold substantial market power and public pharmacies have a cost advantage. However, public pharmacies are of lower quality than their private counterparts: They require consumers to travel more than two times more, carry less product variety, and have more restrictive operating hours and longer waiting times.

To estimate the impacts of public pharmacies, we combine quasi-experimental approaches with a field experiment to study market outcomes and political preferences. The quasi-experiment exploits the staggered entry of public pharmacies across counties. To support this design, we show that the timing of entry was unrelated to baseline differences or pre-trends in local market attributes. Moreover, anecdotal evidence suggests that the timing of entry of public pharmacies depended partly on unexpected delays in the bureaucratic procedure for obtaining sanitary permits. The field experiment consisted of an informational intervention with consumers, which we conducted during the weeks preceding the 2016 local election in counties with public pharmacies. The treatment covered the existence, location, low prices, and low convenience of public pharmacies. We surveyed consumers before the intervention and two months after, collecting data about drug shopping behavior and political participation.

We begin by estimating how the entry of public pharmacies impacted private-sector market outcomes. We exploit the staggered entry of public pharmacies and drug-level data to estimate their impact on private pharmacy prices and sales. Eighteen months after opening, the average

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<sup>1</sup>Chile has relatively high drug prices and high out-of-pocket spending as a share of health expenditures compared with other OECD countries (OECD, 2015).

public pharmacy had shifted 4 percent of sales away from private pharmacies. The decrease in sales was concentrated among drugs that target chronic conditions. We also find a *positive* and growing effect of public pharmacies on private sector prices: By the end of our sample period, the entry of public pharmacies had induced private pharmacies to increase their prices by 1 percent. We interpret this positive price effect as evidence that this low-price and low-quality public option generated market segmentation. In particular, private pharmacies responded to a shift of relatively price-sensitive consumers toward public pharmacies—and thus a less elastic residual demand—by increasing prices. This result is consistent with theoretical research on the potential for price-increasing competition (Chen and Riordan, 2008). A simple model of competition with differentiated firms rationalizes the lack of a stronger demand shift to public pharmacies, despite their low relative prices, as a result of low relative quality. These results show that public pharmacies generated winners and losers as a consequence of their equilibrium effects.

The reduction in consumer drug expenditure generated by public pharmacies compensates for their costs. We develop a simple accounting framework to implement this comparison. First, we estimate the cost of public pharmacies using detailed data on municipal finances. We find that public pharmacies increase net municipal spending on health services by 4.9 percent, and health services revenue by 3.5 percent. Our estimates do not allow us to rule out that this small financial burden came at the cost of foregone increases in net spending on other services. Second, we quantify the benefits public pharmacies provide to consumers. Combining our estimates of economic effects with summary statistics on drug expenditures and prices, we find that introducing public pharmacies in every county would reduce yearly drug expenditure by 1.6 percent or US\$61.5 million, which is 8.7 percent higher than the cost of the policy.<sup>2</sup> Equilibrium price responses by private pharmacies are quantitatively relevant, and omitting them would lead to overestimating the reduction in expenditure by 64 percent.

Budget constraints and electoral incentives are crucial drivers of policy decisions (Besley and Case, 1995; Lizzeri and Persico, 2001; List and Sturm, 2006). Although we document that public pharmacies are relatively low cost and descriptive patterns suggest that mayors expected political returns, their small negative impact on a large number of people suggests that this policy might not be politically profitable. Using our field experiment, we provide suggestive evidence showing that the entry of public pharmacies increased political support for incumbent mayors. In particular, we show that awareness of the availability and attributes of a public pharmacy increased the likelihood of supporting the mayor by 6 percentage points in the local election, although point estimates are

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<sup>2</sup>In addition to its economic effects, increased access to drugs could improve prescription adherence and thus health outcomes. Using data on avoidable hospitalizations and deaths, we find no evidence of such effects. This null result justifies our focus on reduced drug expenditure as a measure of benefits from public pharmacies.

only marginally significant at conventional levels. We combine these results with our estimates of economic effects and we cannot rule out that public pharmacies have a political return that is similar to that of cash transfers (Manacorda et al., 2011).

Overall, we show that public pharmacies created winners and losers: Consumers who switched to public pharmacies benefited from lower prices and, those who did not, lost from higher prices. The public option did not become a financial burden because of its higher bargaining power in the input market and because private firms hold substantial market power in the wholesale and retail markets. Our paper highlights that state-owned firms could be particularly effective in other contexts in which these two conditions are also met. By doing so, we inform the long-standing question of state versus private ownership of firms and the desirability of introducing a public option into otherwise private markets. Access to a public option exists in a variety of settings, including trash collection, mail delivery, housing finance, and internet service providers in the U.S., and historically in retail gasoline stations in Canada (Petro Canada). Recent calls for the introduction of a public option in the U.S. include non-commercial banking, mortgages, and most notably healthcare.<sup>3</sup>

Most previous empirical work has studied public competition in the context of large programs in education (Epple and Romano, 1998; Hoxby, 2000; Dinerstein and Smith, 2021; Dinerstein et al., 2022) and health insurance (Duggan and Scott Morton, 2006; Curto et al., 2019; Saltzman, 2023). Recent work has focused on the role of state-owned firms in local markets, either directly managed by the central government, as in the case of milk stores in Mexico (Jiménez-Hernández and Seira, 2022) and branches of government-owned banks in Brazil (Fonseca and Matray, 2022), or outsourced to the private sector in the Dominican Republic and Indonesia (Busso and Galiani, 2019; Banerjee et al., 2019). Relatedly, Handbury and Moshary (2021) study the price responses of grocery stores following the expansion of the national school program in the U.S. This work mostly finds that prices decrease upon increasing public competition. Our paper contributes to this literature by studying the effects of the entry of locally managed state-owned firms into local pharmaceutical markets, and by showing that public competition can potentially induce market segmentation and lead to an increase in prices by private firms.

This paper also contributes to a literature that studies how store entry affects local market outcomes (Basker, 2007; Hausman, 2007; Jia, 2008; Matsa, 2011; Atkin et al., 2018; Arcidiacono et al., 2020; Bergquist and Dinerstein, 2020). The extent to which entry can generate segmentation in differentiated product oligopoly markets has been studied theoretically by Chen and Riordan

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<sup>3</sup>See, e.g., “Why America needs a public option for mortgages” by Jeff Spross (The Week, 2017), or “There Should Be a Public Option for Everything” by Ganesh Sitaraman and Anne L. Alstott (*New York Times*, 2019).

(2008). Empirically, Frank and Salkever (1997) and Ward et al. (2002) provide evidence for price increases by incumbent products upon the entry of generic drugs and private-label consumer packaged goods. We contribute to this literature by studying the consequences of entry by low-price and low-quality firms and providing evidence of market segmentation.

Our analysis of political support for incumbent mayors who opened public pharmacies is related to a large literature that studies whether and how information about politicians and policies can shape political preferences. Previous research has studied the impact of information on the candidates in an election, incumbent policies, and the prevalence of corruption (Ferraz and Finan, 2008; Gerber et al., 2011; Chong et al., 2015; Kendall et al., 2015; Dias and Ferraz, 2019). Our experimental analysis differs from previous work by providing information on a specific policy directly to the people most likely to be affected by it and only a few weeks before the election.<sup>4</sup> More generally, we contribute to the literature by providing novel evidence of political returns to the introduction of state-owned firms in local markets.

Finally, this paper contributes to the literature that analyzes policies that aim to increase access to pharmaceuticals. Although access to affordable drugs is a first-order policy concern in low- and middle-income countries, which policies regulators should implement to achieve this goal is up for debate (UN, 2010; Pinto et al., 2018). Recent work examines the effects of increased competition in the retail market. Moura and Barros (2020) study the price effects of competition in the market for over-the-counter drugs, while Bennett and Yin (2019) study the price and quality effects of the entry of pharmacy chains in a market dominated by low-quality firms. Other research focuses on the effects of policies to lower drug prices, including price regulation (Dubois and Lasio, 2018; Dubois et al., 2022; Mohapatra and Chatterjee, 2020; Maini and Pammolli, 2022); quality regulation (Atal et al., 2022); and public procurement (Brugués, 2020; Dubois et al., 2021). We provide novel evidence of how public competition in the retail market affects equilibrium market outcomes.

## **1 The Public Option in Retail Pharmaceutical Markets**

Before the introduction of public pharmacies in Chile, consumers could obtain pharmaceutical drugs by buying from private pharmacies or from public health care providers. According to the 2016-2017 National Health Survey (*Encuesta Nacional de Salud*, ENS), almost 40 percent of pharmaceuticals were purchased in the private retail sector, in which there is limited insurance coverage; pharmaceuticals are the most important item of out-of-pocket health expenditures in the

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<sup>4</sup>The focus on health relates our paper to recent work on the effects of the Medicaid Expansion on voter registration and turnout (Haselswerdt, 2017; Clinton and Sances, 2018; Baicker and Finkelstein, 2019).

country (OECD, 2015; Benítez et al., 2018).<sup>5</sup> The private sector is highly deregulated, as there are no market structure regulations or price controls. The three largest chains account for around 80 percent of the market share (FNE, 2019), and stores are geographically clustered in relatively rich areas (MINECON, 2013). Average profit margins in the retail sector reached 40 percent during our period of study (FNE, 2019). The wholesale market is also highly concentrated. According to data from the Economic National Prosecutor (*Fiscalía Nacional Económica*, FNE), 72 percent of off-patent medical products—defined as a unique combination of an active ingredient and a dosage—are produced by only one manufacturer, and 99 percent of those markets have an HHI above 2,500. Moreover, profit margins for manufacturers of off-patent products were 52 percent on average (FNE, 2019).<sup>6</sup>

The rise of public pharmacies was preceded by a collusion scandal in the pharmaceutical industry in 2008 that involved the three largest pharmacy chains in the country (Alé-Chilet, 2018). In a high-profile antitrust case, the pharmacy chains were found guilty. A left-wing mayor of a large county responded to public demands and opened the first public pharmacy in October 2015. Soon after, the popularity of the mayor boomed and dozens of other mayors from all political parties decided to open public pharmacies in the following months. By the end of 2018, 146 out of the 344 counties in the country were operating a public pharmacy. Figure 1 plots the number of counties with a public pharmacy over time, and Figure A.1 displays photos of a private and a public pharmacy.

Public pharmacies offer lower prices because they operate as nonprofit firms by law and have a cost advantage. The latter comes to a large extent from their ability to use a public intermediary that aggregates demand from public providers—most importantly, public hospitals and primary care centers—to negotiate lower prices with manufacturers. As we discuss in detail in Section 2.1 below, around two-thirds of public pharmacies purchase most of their drug supplies through the public intermediary (as opposed to directly from manufacturers). The beneficiaries of public pharmacies are determined by a combination of eligibility requirements, health conditions, and location. Most public pharmacies require that consumers reside in the county, which is determined through a simple enrollment process that entails showing proof of residence. Also, most public pharmacies offer prescription drugs with a focus on drugs that target chronic conditions. Hence, individuals with chronic conditions are more likely to benefit. Finally, public pharmacies enter the

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<sup>5</sup>There is no broad prescription drug insurance market in Chile. Instead, there are a few disjoint programs that mostly cover drugs in the public network or for a limited set of diseases.

<sup>6</sup>Using a broader definition of a market that includes different dosages of the same active ingredient (ATC5), the share of single-firm markets is 54 percent. Still, 89 percent of markets have an HHI above 2,500 under that market definition.



market with a single location per county, whereas there are multiple private pharmacies in each market; this implies that for most consumers, travel costs to public pharmacies are higher than to private pharmacies.

The increasing popularity of public pharmacies has been accompanied by economic and political controversies. On the economic side, there are two main criticisms. First, that public pharmacies may be financially unsustainable and could become a burden for local governments. Second, that public pharmacies could be a form of unfair competition, particularly with respect to non-chain private pharmacies—which accounted for around 20 percent of the market, had limited buying power, and were not involved in the collusion scandal. These criticisms motivate part of our analysis, particularly the impact of public pharmacies on private sector outcomes and municipal finances.

## 2 Research Design

### 2.1 Data

We collected the opening dates and locations of public pharmacies. Openings span the period between October 2015 and April 2018. Figure 1 shows the number of openings per month and the evolution of the total number of public pharmacies operating over time. Their opening before the local election on October 23, 2016—in which most incumbent mayors were running for reelection—seemed far from a coincidence for many. The abrupt increase in openings during the months before the election is hard to explain without resorting to a political argument.

Regarding the supply of drugs by public pharmacies, we exploit detailed data on drug purchases for the 96 pharmacies that have used the public intermediary. These data include the name, molecule, dosage, amount, and price of every drug transaction by public pharmacies in 2016–2018. These data provide information on wholesale (as opposed to retail) prices, but public pharmacies charge low or no markups. While these data cover purchases through the public intermediary in detail, we have only limited data on direct purchases by public pharmacies from manufacturers. Therefore, we are unable to measure aggregate sales by public pharmacies and hence we cannot estimate the impact of their entry on aggregate sales in the market. Our limited data on direct purchases to manufacturers suggest that public pharmacies that deal with the public intermediary purchase most of their drugs through that channel.<sup>7</sup> Hence, we consider that the data from the pub-

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<sup>7</sup>With the goal of measuring the relative relevance of the public intermediary as a supplier of public pharmacies, we collected additional data on public pharmacy direct purchases to manufacturers through data requests. Using data from a sample of 14 counties for which we obtained such information, we estimate that the public intermediary accounts for around 70 percent of total purchases by public pharmacies, and is hence their main supplier. This finding

lic intermediary provides a fairly accurate characterization of public pharmacies. Therefore, we use these data in Section 3.1 to describe how prices, quantities, and variety in public pharmacies compare with those in private pharmacies.

To measure outcomes for private pharmacies, we use data from IQVIA, a company that collects pharmaceutical market information worldwide. These data contain monthly local drug prices and sales for 2014-2018 collected from two sources. The four largest pharmacy chains, which account for more than 90 percent of market share, report retail prices and sales directly to IQVIA. Data for other pharmacies are collected from wholesalers.<sup>8</sup> IQVIA aggregates the data at the level of 66 local markets, which cover most of the country.<sup>9</sup> We restrict our attention to prescription drugs, which account for 93 percent of the drugs among the molecules we include in the analysis.

## 2.2 *The Entry of Public Pharmacies*

In this section, we describe entry patterns of public pharmacies and discuss how they can be exploited to study their effects. We begin with a characterization of the counties that opened a public pharmacy. We then study the timing of the entry of public pharmacies and their location within the counties in which they opened. Our results show that counties that open public pharmacies differ systematically from those that do not, but the timing of opening among those that open does not seem to be driven by observable county characteristics.

We start by comparing counties with and without public pharmacies. Columns (1)-(3) in Table 1 show these results. Panels A and B show that public pharmacies opened in dense high-income counties with more penetration of private health insurance, slightly better self-reported health, and a private pharmaceutical market with more pharmacies, more sales, and higher prices. In contrast, Panel C shows few differences in political variables, as measured by the previous local election of 2012.<sup>10</sup> If anything, counties with a public pharmacy had more candidates and were more likely to have a winner from the left wing. In sum, counties with and without public pharmacies differed

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motivates using the detailed data from the public intermediary in order to describe the attributes of public pharmacies.

<sup>8</sup>We adjust these prices for inflation using the health CPI from the National Institute of Statistics and compute prices per gram of the active ingredient to normalize them across presentations.

<sup>9</sup>Moreover, the data provide price and sales information at the product level for branded drugs, which identifies the laboratory, dosage, and presentation of each drug. However, for unbranded drugs it only provides price and sales information at the dosage and the presentation level, aggregated across laboratories. This is irrelevant for our analysis since we focus on price indices and aggregate sales at the molecule level.

<sup>10</sup>In Chile, all mayors are elected simultaneously by a simple majority rule in elections held every four years and without term limits until 2020. To measure local political outcomes, we use county-level information about candidates, parties, coalitions, and votes for each candidate in the 2012 and 2016 local elections from the Electoral Service. The 2012 election allows us to characterize the political equilibrium before the opening of public pharmacies.

significantly in terms of their pharmaceutical market and socioeconomic characteristics but were relatively more similar in their political characteristics.

To examine entry timing systematically, we ranked all public pharmacies by their entry date and estimated an ordered logit model of this ranking on all variables in Table 1. Column (4) in the table presents the results. Pharmacies that opened earlier entered counties with more population and were more likely to have left-wing mayors, but entry timing is otherwise uncorrelated with the characteristics of the pharmaceutical market, socioeconomic attributes, or electoral competition in the previous election. Instead, anecdotal evidence suggests that unexpected delays in sanitary permits explain why some pharmacies opened after the election. We rely on these results to exploit the timing of entry as exogenous variation.

Finally, we document that mayors opened public pharmacies near existing private pharmacies, which provides a unique opportunity to study the impact of the public option in an existing market. To describe their location choices, we geocoded all private pharmacies in the country and assigned them to geographic cells of 600×600 meters. We then estimated cross-sectional cell-level regressions using data from counties with a public pharmacy. The dependent variable is an indicator for a cell that has a public pharmacy, and explanatory variables include the number of private pharmacies, the number of schools as a proxy for population, and county-level fixed effects. Table A.1 displays the results. Estimates reveal that public pharmacies opened in populated areas where private pharmacies were already operating. The maps in Figure 2 provide visual examples of the entry decision in six counties spread across the country.

### **3 The Economic Effects of Public Pharmacies**

#### *3.1 Evidence on Prices and Quality of Public Pharmacies*

When public pharmacies opened, consumers gained access to a new alternative in their choice set, which differed from available options along several dimensions. We describe the basic attributes of public pharmacies by using transaction-level data on all purchases by public pharmacies from the public intermediary in 2016–2018. The public intermediary was the main supplier of drugs for the 96 counties that sourced through it, as discussed in Section 2.1.

The most salient and advertised difference was related to drug prices. Using a set of exactly matched drugs that are sold in both public and private pharmacies, we study price differences across public and private pharmacies. In Figure 3, Panel (a) shows that almost all drugs are sold at lower prices in the former and that the relative price difference is, on average, between 64 and 68 percent depending on the margin public pharmacies charge over purchase costs from the public

intermediary. These large price differences suggest that consumers should, in principle, switch to public pharmacies in the local markets in which they open.

Two leading reasons for these price differences are public pharmacies' higher bargaining power in the input market—coupled with a concentrated input market—and the substantial market power of private retailers downstream, both of which we discussed in Section 1. In Online Appendix A, we formalize these arguments by developing a model of the vertical chain that captures the main features of our setting—namely, that (i) producers and retailers are able to exercise market power, (ii) state-owned firms differ from private firms by having greater bargaining power upstream, and (iii) state-owned firms do not maximize profits but rather total surplus. We show that under mild assumptions regarding the demand curve, downstream prices are lower when retailers have more bargaining power upstream and when retailers place a higher weight on consumer welfare relative to profits.

Consumers trade off lower prices with the lower quality of public pharmacies. The fact that public pharmacies enter with a single store in each county implies that most consumers have multiple private pharmacies closer to their homes. Using data on voter home addresses from the Electoral Registry, and the locations of public and private pharmacies, we calculate distances between households and every pharmacy in the county. The average (median) individual has 20 (12) private pharmacies located closer than the public pharmacy in their county. Panel (b) in Figure 3 shows that the distributions of distance to the closest private pharmacy and public pharmacy differ markedly: The average distance to the closest private pharmacy is 1.1 kilometers—less than half of that to the public pharmacy. These facts imply that shopping at public pharmacies entails higher travel costs than shopping at private pharmacies. Moreover, public pharmacies offer less product variety. Panel (c) in Figure 3 shows that the average number of products per molecule-county is 2.2, and that 70 percent of molecule-counties offer 3 varieties or fewer, while the average number of varieties in private pharmacies is 15.2.<sup>11</sup> To the extent that consumers value product variety, these patterns imply that public pharmacies are less convenient than private pharmacies. The longer waiting times and limited opening hours already described in Section 1 further exacerbate the relatively low quality of public pharmacies.

The relevance of public pharmacies has grown over time, which demonstrates that at least some consumers value lower drug prices relative to lower convenience enough to switch to public pharmacies. Panel (d) in Figure 3 shows that their average market share across molecules and counties reached around 4 percent by the end of 2018. Of course, it is unclear whether sales by

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<sup>11</sup>Relatedly, public pharmacies are more likely to offer only generic drugs or only branded drugs within a molecule: This is the case for 72 percent of molecule-counties at public pharmacies, but for only 36 percent at private pharmacies.

public pharmacies have decreased sales by private pharmacies or simply expanded market size. To inform this margin, we estimate the effects of public pharmacies on private pharmacy sales.

### 3.2 *Equilibrium Effects on Prices and Sales by Private Pharmacies*

Public pharmacies may induce consumers to substitute away from private pharmacies.<sup>12</sup> Moreover, the competitive pressure from public pharmacies may induce private pharmacies to adjust prices. In this section, we estimate the effects of the entry of public pharmacies on prices and sales by private pharmacies.

Theoretically, the effects of entry on incumbent firm prices are ambiguous. Chen and Riordan (2008) study the conditions under which entry leads to increases or decreases in prices. Their analysis shows that these effects depend on the magnitudes of two effects of entry on the incumbent's pricing incentives. First, entry has a *market share effect*, which depends on the extent to which the incumbent loses demand upon entry due to substitution. The more demand the entrant takes away from the incumbent, the stronger the incentives for the incumbent to decrease prices in response to entry. Second, entry has a *price sensitivity effect*, which depends on how the slope of the incumbent's residual demand curve changes after entry. The steeper the demand curve after entry relative to before entry, the lower the extent of substitution away from the incumbent upon entry, and therefore the stronger its incentive to increase prices upon entry. Overall, the incumbent's price will increase whenever the price sensitivity effect dominates the market share effect and vice versa. Which effect dominates depends on the distribution of consumer preferences and on the attributes of the firms. To further develop intuition for the conditions under which private pharmacy prices may decrease or increase upon the entry of public pharmacies, we develop a model based on Chen and Riordan (2008) in Online Appendix C. We then implement illustrative simulations that we employ to discuss our results.

**Event study evidence.** We start by exploiting the staggered entry of public pharmacies in an event study framework. For this analysis, we use IQVIA data on drug prices and sales across local markets. A challenge in combining data on the entry of public pharmacies with data from IQVIA is that the level of geographic aggregation of the latter markets is in some cases larger than

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<sup>12</sup>As part of this research, we designed and implemented an informational field experiment to study the impacts of public pharmacies. In the experiment, we randomly provided information about public pharmacies to individuals buying pharmaceuticals in private pharmacies. In this paper, we use the experiment to estimate the impact of public pharmacies on support for incumbent mayors who opened these. We provide more details in Section 5. However, we also collected data on consumer shopping behavior both before and two months after the intervention, to study whether consumers in the pharmaceutical market switched from private to public pharmacies. Overall, consumers learned about the low-price and low-quality of public pharmacies after the intervention, and to some extent reported either having used or planning to use the public pharmacy. We discuss these findings in Online Appendix B.

counties, which is the level at which public pharmacies operate. To tackle this issue, we estimate a stacked event study regression.<sup>13</sup> Whenever a market has more than one event, we create as many copies of the data as the number of events. We stack the copies in a dataset and use the entry of public pharmacies to all counties within a market as events. Figure A.2 shows the distribution of the number of events per market.

The main specification we estimate is given by:

$$y_{mlgt} = \sum_{k=-12}^{18} \beta_k D_{lgt}^k + \lambda_{mt} + \theta_{mlg} + \varepsilon_{mlgt}, \quad (1)$$

where  $g$  indexes entry events within a market. The dependent variable  $y_{mlgt}$  is either the log of drug prices or the log of drug sales for molecule  $m$  in local market  $l$  in month  $t$ .<sup>14</sup> Our interest is in the coefficients  $\beta_k$  on the dummies  $D_{lgt}^k = 1\{t = e_{lg} + k\}$ , which indicate whether a month  $t$  is exactly  $k$  months after event time  $e_{lg}$  for event  $g$  in local market  $l$ . We normalize  $\beta_{k=-1} = 0$ , so we interpret all coefficients  $\beta_k$  as the effect of a public pharmacy’s opening on the dependent variable exactly  $k$  months after its entry. The specification also includes molecule-month fixed effects  $\lambda_{mt}$  to account for time-varying unobservables at the level of molecules, and molecule-market-event fixed effects  $\theta_{mlg}$  to account for persistent differences in market conditions across markets. Standard errors are clustered at molecule-market level.<sup>15</sup>

The entry of public pharmacies had meaningful effects on private pharmacies. Panels (a) and (b) in Figure 4 present the results for sales and prices, respectively. Drug sales by private pharmacies decrease after a public pharmacy enters a market. Our estimates imply that 18 months after the entry of a public pharmacy, private pharmacies in that market sell around 4 percent less. Furthermore, 18 months after the entry of a public pharmacy, drug prices in private pharmacies increase by

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<sup>13</sup>This approach has been adopted in recent work that estimates event study models in settings with multiple events per unit (see, e.g., Lafortune et al. 2018; Cengiz et al. 2019).

<sup>14</sup>We define the market-level price as the share-weighted average of log prices:

$$\hat{P}_{mlt} = \sum_{i \in I_{ml}} w_{i|0} P_{ilt},$$

where  $I_{ml}$  is the set of drugs of molecule  $m$  in local market  $l$ ,  $P_{ilt}$  is the log price per gram of product  $i$  in period  $t$  and market  $l$ , and  $w_{i|0}$  denotes the share of sales of drug  $i$  in market  $l$  in 2014. Because these weights are constant, changes in the index are driven by changes in prices and not by changes in market shares or market structure. This price index has been used in previous work studying retail drug pricing (e.g., Atal et al., 2022). For sales, we use the residuals from the projection of the outcome variable on month-of-the-year fixed effects by molecule-market to account for seasonality that is specific to sales in some markets (e.g., due to tourism in the summer).

<sup>15</sup>We use a balanced sample of markets in event time and include never-treated markets to pin down the linear component of pre-trends (Borusyak et al., 2022). Moreover, we fully saturate the model and report results for event dummies 12 months before and 18 months after the event.

1 percent. Both effects increase over time, which suggests that public pharmacies evolve in terms of enrolling more consumers and possibly improving their product offerings and convenience.<sup>16</sup>

The main threat to identification of the effect of public pharmacies is reverse causality; unobserved determinants of sales and prices in the private sector may drive the entry of public pharmacies. In that case,  $\beta_k$  would confound the causal effect of public pharmacies on private market outcomes with trends in outcomes that cause the entry of public pharmacies.<sup>17</sup> Reassuringly, the lack of pre-trends in both sales and prices leading up to the entry of public pharmacies suggests that reverse causality and strategic considerations do not play a significant role in our setting.<sup>18</sup>

Another concern relates to multiple public pharmacy entries within a market, which could potentially turn the treatment effect of a previous public pharmacy entry into a pre-trend for the subsequent entry. This concern is muted in our context because the majority of markets experience 1 or 2 events and most subsequent entry occurs within 1 or 2 months of each other, as shown by Figure A.2. To assess the importance of this issue in our setting, we do two robustness checks. First, we redefine the event as the *first* entry of a public pharmacy, in which case this type of pre-trend is absent by definition. The results under that treatment definition are essentially the same as those in our main specification, as shown by Figure A.3. Second, we restrict the estimating sample to markets with a single event or multiple events separated by less than 1 month. The results for this sample track closely those from our main sample, as shown by Figure A.4.

**Exposure difference-in-differences design.** We complement the event study design with a regression analysis that relates market-level outcomes to the share of the population in each market that has access to a public pharmacy at each point in time. The advantage of this design is that it exploits all the variation in the timing of entry of public pharmacies as well as the heterogeneous exposure of markets to public pharmacies. We then employ this design to develop a heterogeneity analysis for the effects of public pharmacies.

We define treatment intensity  $E_{lt}$  as the share of the population in market  $l$  with access to a public pharmacy at time  $t$ , and estimate the following specification:

$$y_{mlt} = \lambda_{mt} + \theta_{ml} + \beta^{\text{jump}} E_{lt} + \beta^{\text{phase in}} E_{lt}(t - t_e^* + 1) + \varepsilon_{mlt}, \quad (2)$$

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<sup>16</sup>An additional margin of response for private pharmacies would be to adjust product variety. We estimate equation (2) using the number of varieties offered as the dependent variable, and find no evidence of responses along that margin.

<sup>17</sup>Strategic entry is an identification threat for reduced-form models for the effects of firm entry as equation (1), but it is not a relevant concern in our context. Public pharmacies' business model differs from private pharmacies' since they operate as nonprofit firms.

<sup>18</sup>As an additional piece of supporting evidence, in column (4) of Table 1 we study the order of entry of public pharmacies using an ordered logit regression of entry on market and political covariates. The results show that the timing of entry is uncorrelated with covariates associated with the supply and demand of drugs.

where  $E_{it} = 0 \forall t < t_e^*$ . This functional form is motivated by the patterns of the treatment effects we estimate in our event study analysis in Figure 4. The parameter  $\beta^{\text{jump}}$  is a mean shift in outcome  $y_{mt}$  after the adoption of a public pharmacy. Since results from the event study specification imply that the impact on sales and prices evolves over time, we allow for a trend break,  $\beta^{\text{phase in}}$ . We include event-time dummies as controls for all periods before  $k = -12$  and after  $k = 18$  in treated markets, for comparability with the event study results. Our main parameter of interest is the effect of the public pharmacy 18 months after its entry, which we calculate as  $\bar{E}_{18} \times [\beta^{\text{jump}} + (18 + 1)\beta^{\text{phase in}}]$ . The term  $\bar{E}_{18}$  is the average exposure to a public pharmacy across markets 18 months after the entry of the first pharmacy in the market.

For ease of exposition, we present the results of the main parameter of interest in Table 2 and report the underlying estimates  $\beta^{\text{jump}}$  and  $\beta^{\text{phase in}}$  in Table A.2. Columns (1) and (2) in Table 2 present estimates for sales and prices, respectively. Panel A shows that the entry of public pharmacies decreases drug sales by private pharmacies by 3.8 percent and increases drug prices by private pharmacies by 1 percent 18 months after their introduction. Reassuringly, these magnitudes are close to the estimates we obtain at the end of the time window in the event studies in Figure 4. To put the magnitude of this estimate in context, the average coefficient of variation of drug prices across drugs and local markets is 0.08. Hence, our estimates imply that drug prices at private pharmacy prices increase by around 12.5 percent of a (relative) standard deviation after the entry of a public pharmacy.<sup>19</sup>

**Heterogeneity analysis.** The remaining panels in Table 2 present a heterogeneity analysis. The characteristics of the context motivated us to focus on three margins. First, public pharmacies specialize in selling drugs for chronic conditions and thus we expect a larger impact on these drugs. Column (1) in Panel B shows that sales of chronic drugs decrease by 4.5 percent, which is 61 percent more than the 2.8 percent decrease in non-chronic drugs ( $p\text{-value} < 0.01$ ).<sup>20</sup> In contrast, column (2) in Panel B shows similar price increases for both types of molecules. Second, we have emphasized quality differences across public and private pharmacies. We proxy relative quality by the ratio of drug variety within each molecule in public pharmacies relative to private pharmacies

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<sup>19</sup>The extent of price variation in our data is somewhat higher than roughly comparable measures for within-chain pricing reported by Adams and Williams (2019) and DellaVigna and Gentzkow (2019) for construction materials and consumer-packaged goods in the U.S, respectively. This price variation is consistent with our ability to estimate price effects in this setting. Results available from the authors.

<sup>20</sup>We observe 102 chronic molecules and 74 non-chronic molecules. This finding is consistent with our experimental evidence showing that households with members with chronic conditions react more strongly to the availability of public pharmacies in terms of shopping behavior. We discuss experimental results in Online Appendix B.



in each market.<sup>21</sup> Column (1) in Panel C shows that the impact is larger in markets in which the public pharmacy has a richer variety of products within each molecule ( $p$ -value 0.02). Column (2) in Panel C reveals larger price responses in markets in which public pharmacies offer less variety of products within a molecule ( $p$ -value<0.01). Finally, we consider whether the spatial distribution of private pharmacies matters for the impacts of public pharmacies. We expect that the closer public pharmacies locate to private pharmacies, the larger the decrease in private pharmacy sales. Column (1) in Panel D presents heterogeneous effects along this dimension and confirms this intuition ( $p$ -value<0.01).<sup>22</sup>

### 3.3 Discussion

The entry of public pharmacies had equilibrium effects on private pharmacies. As expected, due to the lower prices offered by public pharmacies, some consumers substituted away from private pharmacies and drug sales in the latter decreased. While increased competition could have induced private pharmacies to reduce drug prices, we find that private pharmacies instead increased prices. This response is consistent with the price sensitivity effect of entry dominating the market share effect of entry. In particular, while some consumers switched to public pharmacies upon their entry, it must be that they had a relatively low willingness to pay for private pharmacies, which led to the residual demand for private pharmacies to become steeper. The increase in private pharmacy prices we estimate implies that the upward pricing pressure from the latter was larger than the downward pricing pressure from overall substitution toward public pharmacies.<sup>23,24</sup>

The sales response to the entry of public pharmacies may seem small, given the magnitude

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<sup>21</sup>We define high (low) variety as observations above (below) the median of the ratio between the number of distinct products within molecule and market offered by the public pharmacy and those by private pharmacies.

<sup>22</sup>To split the sample in two, we use the average number of public pharmacies operating within 400 meters of private pharmacies. For consistency, we only consider private pharmacies that appear in our data for private pharmacy outcomes. These results need to be interpreted with caution as public pharmacies mostly locate nearby private pharmacies and information about how distance affects pharmacy choice is lacking.

<sup>23</sup>In our model in Online Appendix C, we show that a key condition under which private pharmacy prices are more likely to increase is a negative correlation in consumer willingness to pay for public and private pharmacies, such that consumers who have a high valuation for private pharmacies also have a low valuation for public pharmacies. This negative correlation implies that consumers who substitute away from the private pharmacy upon entry are those with low willingness to pay for the private pharmacy—and thus the most price sensitive—which leads to the residual demand curve of the public pharmacy’s being steeper after entry. In addition, there must be enough heterogeneity in willingness to pay across consumers, as otherwise there is no scope for increasing prices substantially. Figure A.5 shows simulation results that demonstrate that the direction of the price effects of entry indeed depends on these parameters of the distribution of consumer preferences.

<sup>24</sup>Caves et al. (1991) and Frank and Salkever (1997) document a similar pattern of market segmentation in pharmaceuticals, in which innovator drugs that become off-patent do not decrease but rather *increase* their prices after generic entry. This fact is known in the literature on competition in pharmaceutical markets as the “generic paradox.”

of the price differences between public and private pharmacies. Our interpretation is that product differentiation plays a role in mediating this response. As documented above, public pharmacies are less convenient than private pharmacies in terms of waiting times, opening hours, product variety, and travel distance. The lack of a stronger response suggests that a sizable share of consumers value those attributes enough to not substitute toward public pharmacies on the basis of lower prices. Higher-quality public pharmacies would have likely led to stronger equilibrium responses.<sup>25</sup> Second, our event study results in Figure 4 show that both quantity and price effects increase over time, which suggests that the full effects may be larger once the market settles into a new equilibrium.

The substitution away from private pharmacies we estimate is consistent with findings in related work by Busso and Galiani (2019) and Jiménez-Hernández and Seira (2022) in different contexts. However, they find a price decrease among private firms as opposed to a price increase. Our results highlight the fact that the price effects of public competition will depend on underlying consumer preferences and firm attributes.

## 4 The Benefits and Costs of Public Pharmacies

This section discusses the relative efficiency of state-owned firms. First, we estimate the cost of public pharmacies by exploiting data on municipal finance to study the effects of introducing public pharmacies on spending and revenue on health and non-health services. Second, we assess whether public pharmacies have any health effects on consumers as measured by avoidable hospitalizations. Finally, we develop a simple framework that exploits our estimates of the price and quantity effects of public pharmacies to estimate how consumer drug expenditure decreases as a result of public pharmacies, and compare it with our cost estimates.

### 4.1 *Municipal Finance and the Cost of Public Pharmacies*

Given that public pharmacies were created by local governments that manage multiple other local services, it is important to identify whether they are economically sustainable or represent a financial burden that may crowd out other services. To study this margin, we exploit administrative data

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<sup>25</sup>We illustrate the role of vertical differentiation between private and public pharmacies using our model in Online Appendix C. Our model simulations show that vertical differentiation indeed influences the extent to which the entry of public pharmacies affects private pharmacy prices, and market share depends on vertical differentiation. Panel A in Figure A.6 shows that the extent of business stealing by an entrant decreases substantially as the quality of the entrant relative to the incumbent decreases. Moreover, Panel B in Figure A.6 shows that the incumbent in the market is able to sustain higher prices when the quality of the entrant relative to the incumbent is lower.

from municipal finances to estimate the financial impacts of public pharmacies.<sup>26</sup>

For this analysis, we estimate the following regression:

$$y_{ct} = \theta_c + \lambda_t + \pi^{\text{jump}} PP_{ct} + \pi^{\text{phase in}} PP_{ct}(t - t_e^* + 1) + \varepsilon_{ct}, \quad (3)$$

where  $y_{ct}$  is a financial outcome in county  $c$  and year  $t$  (e.g., spending on health services),  $PP_{ct}$  indicates the share of the year with a public pharmacy in county  $c$ , and  $t - t_e^*$  measures the number of years since the opening of the public pharmacy. The specification includes county fixed effects  $\theta_c$  and year fixed effects  $\lambda_t$ . Similar to our specification for private market outcomes in equation (2), the parameter  $\pi^{\text{jump}}$  captures a mean shift in the dependent variable after treatment, whereas  $\pi^{\text{phase in}}$  captures a trend break. In terms of data, we observe annual county spending and revenue for 2013–2019. Both spending and revenue have accounts that we aggregate into health and non-health categories. Moreover, within the accounts related to health spending and revenue, we construct measures of spending by and revenue from public pharmacies. To ease comparisons across counties, we use the log spending and revenue per capita as dependent variables in this analysis.<sup>27</sup>

Table 3 presents our main results and Table A.3 presents coefficient estimates of equation (3). The main result is the effect of public pharmacies after 18 months of operation (1.5 years), which we compute as  $\pi^{\text{jump}} + (1.5 + 1) \times \pi^{\text{phase in}}$ . The results deliver four main messages. First, 18 months after the entry of public pharmacies, we observe an increase of 30 percent on health spending related to the public pharmacy, along with a 20.6 percent increase in revenue related to the public pharmacy. Second, these impacts are also statistically significant when looking at health spending and revenue more broadly: We estimate an increase of 4.9 percent in health spending 18 months after the entry of public pharmacies in column (3), which is partially compensated for by an increase in health revenue of 3.5 percent in column (4). Third, the impacts of public pharmacies on non-health services in columns (5) and (6) are imprecisely estimated and we cannot rule out a decrease of a magnitude similar to the increase in health services. Fourth, in terms of overall municipal finance, our point estimates in columns (7) and (8) imply that spending increases more than revenue, although those coefficients are again not statistically significant. Taken together, the point estimates in the last two columns suggest that public pharmacies induced, if any, only a small

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<sup>26</sup>The data come from the National System of Municipal Information (*Sistema Nacional de Información Municipal*, SINIM). Counties spend resources on transportation, public education, public health, culture, and sports, among others (Law 18695). Approximately 90 percent of their budget comes from county revenues (property and vehicle tax receipts) and other resources correspond to monetary transfers from the central government.

<sup>27</sup>Some counties, which account for 7 percent of the sample, do not report the breakdown of their accounts for health and non-health services. To obtain a uniform sample across dependent variables, we drop those observations.

and statistically insignificant increase in the overall municipal deficit.<sup>28,29</sup>

These estimates allow us to compute the average cost of introducing a public pharmacy. A public pharmacy's profits depend on the markup they charge on drugs if any, and any initial investment and operating cost it incurs. The fact that public pharmacies induce a deficit implies that they set prices below average cost. The average spending and revenue per capita are \$695.68 and \$730.15 and the average county in the country has a population of 51,781. Combining these statistics with our point estimates in columns (7) and (8) of Table 3, we calculate that after 18 months of operation the annual loss for a public pharmacy in the average county is \$164,442.<sup>30</sup> The next sections compare this cost estimate with the estimated benefits of public pharmacies for consumers.

#### 4.2 *Lack of Health Effects of Public Pharmacies*

Increased access to pharmaceutical drugs could benefit individuals through health improvements. For instance, such effects could operate through improved adherence to prescription drugs for individuals with chronic diseases due to lower prices and increased access (Cutler and Everett, 2010). However, in our setting we do not observe individual-level prescriptions and drug purchases. Instead, we focus on avoidable hospitalizations associated with chronic diseases, which would likely have not occurred under appropriate disease management. This variable has been employed previously in the literature (e.g., Layton et al., 2019). The fact that public pharmacies were oriented toward individuals with chronic diseases makes this variable particularly suitable. We would interpret a decrease in avoidable hospitalizations after the entry of a public pharmacy as a signal that the pharmacy increased drug access and, in consequence, adherence by individuals with chronic diseases.

For this analysis, we estimate equation (3) using avoidable hospitalizations as the dependent variable. We exploit data on monthly hospitalizations for 2013–2019 from the Ministry of Health (DEIS, 2019), which cover the number of hospitalizations, days of hospitalization, number of surgeries, and number of deaths per diagnosis across all hospitals in the country. The number of

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<sup>28</sup>Figure A.7 displays corresponding event study estimates and provides reassuring evidence regarding the trends in these outcomes leading up to the entry of public pharmacies.

<sup>29</sup>The data on municipal finance have some zeros, which implies that by taking the log of the dependent variable we drop some observations. This share is not higher than 2 percent across outcomes, so the impact of this transformation is small. Table A.4 shows results from the same specification for alternative transformations of the dependent variable. The main takeaway from this robustness check is that our results are essentially unchanged across these transformations.

<sup>30</sup>Articles from local newspapers that disclose public pharmacy non-drug costs place the yearly cost of running them at between \$85,000 and \$125,000, which likely provide a lower bound for total operating costs and are in line with our estimates (see, e.g., Araucanía Cuenta 2016; El Austral 2017; Clave9 2017; Diario Concepción 2017; Diario Financiero 2022).

hospitalizations captures only the volume of these events, whereas hospitalization days, surgeries, and deaths capture their severity. To focus on the subset of diagnoses for which hospitalizations are considered avoidable, we follow the Prevention Quality Indicators in AHRQ (2019), which lists all diagnosis codes (ICD-10) for avoidable admissions associated with asthma, chronic obstructive pulmonary disease, diabetes, and hypertension. We restrict our sample of hospitalizations for this analysis to these diagnoses. We normalize these variables by population and measure them per 100,000 inhabitants.

Our estimates suggest that public pharmacies did not improve health outcomes, at least in the short period of time we are able to examine. Table 4 presents our main results and Table A.5 presents coefficient estimates of equation (3). For each outcome, we show results for all individuals and for those under public insurance (*Fondo Nacional de Salud*, FONASA), who on average have lower income and are more likely to benefit from a public pharmacy. Across all outcomes and samples, we find no statistically significant effect of the entry of a public pharmacy to a local market after 18 months. That said, our estimates are not precise enough to rule out effects that could be quantitatively meaningful. In particular, our estimates can reject at the 5 percent level reductions of 2.43 hospitalizations, 21.15 hospitalization days, 0.23 surgeries, and 0.07 deaths per 100,000 inhabitants as the effect of public pharmacies, which are equivalent to reductions of between 10 percent and 13 percent in these outcomes relative to their baseline levels.<sup>31,32</sup>

Overall, our interpretation of these results is that public pharmacies did not affect access to drugs to an extent such that adherence improved enough as to reduce avoidable hospitalizations. It is important to note that the lack of a health effect is likely to be mediated by contextual factors such as the elasticity of demand and access to health services, among others. Regardless, these results suggest that if public pharmacies had any market-creation effect, it was small, and most of the effect was through business stealing from private pharmacies.

### 4.3 Comparing Costs and Benefits

In this section, we use our previous results to compare the benefits and costs of public pharmacies. Our measure of benefits from public pharmacies focuses on reduced expenditure in drugs for consumers, given that we find no evidence of health effects. We develop a simple accounting framework to estimate effects on consumer expenditure by combining our results on economic

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<sup>31</sup>Figure A.8 shows the results of an event study version of equation (3). For all outcomes and samples, we again find no evidence that public pharmacies affected health outcomes. Reassuringly, these results show a lack of differential trends across counties leading up to the entry of public pharmacies, which provides evidence against reverse causality.

<sup>32</sup>An additional analysis of school attendance and sick leaves—arguably related to the health of children and the working population—also suggests a null impact of public pharmacies in the short run. See Table A.6 and Figure A.9.

effects from Section 3 with basic statistics from the market.

Let  $r$  denote private pharmacies and  $u$  denote the public pharmacy. Moreover, let  $t = 0$  indicate the period before entry of the public pharmacy and  $t = 1$  the period after its entry. Using this notation, total consumer expenditure in period  $t$  is given by  $e_t = M_t(s_t^r p_t^r + s_t^u p_t^u)$ , where  $M_t$  is the amount of drugs consumers need;  $s_t^r$  and  $s_t^u$  are market shares of the private and the public pharmacy, respectively; and  $p_t^r$  and  $p_t^u$  are composite drug prices at each of them. We impose two assumptions. First, we assume that the market size remains constant over time, such that  $M_t = M$  for  $t = 0, 1$ . Second, given that we are unable to estimate aggregate effects on drug quantity with the available data, we rule out such effects and impose  $s_t^r + s_t^u = 1$  for  $t = 0, 1$ .

The object of interest is the change in drug expenditure upon entry of the public pharmacy:

$$\Delta e = M(s_1^r p_1^r + s_1^u p_1^u) - M(s_0^r p_0^r + s_0^u p_0^u),$$

which we can rearrange to be a function of our estimates and data. First, note that  $s_0^r = 1$  and  $s_0^u = 0$  by definition. Second, we use our estimates of effects on private pharmacies from Section 3.2 to express the sales and prices of private pharmacies after the entry of the public pharmacy as  $s_r^1 = (1 - \beta_s)s_0^r$  and  $p_r^1 = (1 + \beta_p)p_0^r$ , respectively. Finally, we use results from Section 3.1 on price differences between public and private pharmacies to express public pharmacy prices as  $p_u^1 = \phi_1^u p_r^1$ , where  $\phi_1^u$  is the average discount public pharmacies offer relative to private pharmacies. After replacing and rearranging, we get:

$$\Delta e = \underbrace{Mp_0^r}_{\text{Baseline expenditure}} \times \left[ \underbrace{(1 - \beta_s)(1 + \beta_p) - 1}_{\Delta \text{ expenditure in private pharmacies}} + \underbrace{\beta_s \phi_1^u (1 + \beta_p)}_{\Delta \text{ expenditure in public pharmacy}} \right].$$

To measure the change in drug expenditure, we proceed as follows. We measure baseline expenditure using data from the 2016 National Household Spending Survey (*Encuesta de Presupuestos Familiares* EPF) which states that the average yearly drug expenditures were \$213.4. Furthermore, our estimates from Section 3.2 imply that  $\beta_s = 0.039$  and  $\beta_p = 0.010$ . Finally, we know from Section 3.1 that public pharmacies set prices at an average of  $\phi_1^u = 0.34$  of private pharmacy prices.

The average consumer saves \$3.5 per year, according to these estimates. This average masks substantial heterogeneity: Those who stayed at private pharmacies increased their annual spending by \$2.2, whereas those who switched to the public pharmacy reduced theirs by \$143.8. A population of particular interest is consumers with chronic conditions, who are the main target of public pharmacies and account for 22 percent of the population, according to the 2016–2017

ENS. Our estimates imply that these consumers decreased their yearly expenditure by an average of \$17.3. Of them, those who stayed with private pharmacies increased their yearly expenditure by \$8.4, whereas those who switched decreased it by \$551.3. To put these numbers in context, the median monthly wage among working-age individuals in 2017 was around \$670. Adding across consumers, these estimates imply that consumers in the average county decreased their aggregate spending by \$178,777 per year. If all counties in the country introduced public pharmacies, aggregate spending would decrease by \$61.5 million per year—equivalent to 1.58 percent of total expenditure according to the EPF. Accounting for equilibrium price responses by private pharmacies is quantitatively relevant; omitting them would lead to overestimating the reduction in expenditure by 62 percent.

Our estimates imply that consumer benefits in terms of reduced drug expenditure on inframarginal units are 8.7 percent higher than the cost of public pharmacies a year and a half after their entry. Public pharmacies achieve reductions in consumer expenditure higher than their costs for two reasons: public pharmacies hold a cost advantage relative to private pharmacies when purchasing from manufacturers, and private pharmacies hold substantial market power in the retail market (FNE, 2019). Public pharmacies thus address two salient market failures in this industry. Because of this, the introduction of a state-owned firm likely performs better than an alternative policy of subsidizing drug purchases. In this simple framework, the cost of a subsidy is the reduction in drug expenditure, and is thus higher than that of the public pharmacy, according to our estimates. This is because subsidies are able to reduce drug expenditure, but do not address market power in the private market and therefore must incur a higher cost to achieve the same effects as the public pharmacy.<sup>33</sup>

Of course, this is not a full welfare analysis. On the one hand, we do not account for potential market expansion effects, which implies that we may underestimate the benefits of public pharmacies. On the other hand, we do not account for consumer valuation of the relative convenience of private and public pharmacies. The fact that relatively few consumers switch despite the large potential savings for switchers suggests that the valuation of these non-price pharmacy attributes is high.<sup>34</sup> A richer model of consumer demand and pharmacy pricing is needed to conduct such an

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<sup>33</sup>Enriching the framework to account for aggregate effects would exacerbate the extent to which state-owned firms outperform subsidies since subsidies would in that case induce an additional deadweight loss.

<sup>34</sup>To provide a lower bound on the relative inconvenience of public pharmacies, we estimated the cost of additional travel time to public pharmacies. To do so, we combined standard assumptions from the transportation literature with data on (i) the spatial distribution of households, private pharmacies, and public pharmacies, and (ii) the distribution of hourly wages. We find that an individual with an average hourly wage has an average annual cost of additional travel time to public pharmacies of \$13.9, with 25<sup>th</sup> and 75<sup>th</sup> percentiles of \$2.3 and \$21, which are well below our estimates of average savings for switchers. These patterns suggest that while their inconvenient locations may indeed

analysis.<sup>35</sup>

## 5 Political Returns of Public Pharmacies

Budget constraints and electoral incentives are crucial drivers of policy decisions (Besley and Case, 1995; Lizzeri and Persico, 2001; List and Sturm, 2006). The small negative impact on a large number of consumers suggests that the public option might not be politically profitable. This section uses an informational field experiment, along with self-reported voting behavior, to estimate the causal effect of the awareness of public pharmacies among consumers in the pharmaceutical market on political support for the incumbent who opened the pharmacy.

### 5.1 *The Field Experiment*

We designed a field experiment to study whether the availability of public pharmacies affected consumers. To induce variation in awareness of the public pharmacy within local markets, we implemented an informational intervention. The decision to provide information was based on a survey we conducted before the experiment, which revealed that consumers were only partially informed along two dimensions. First, some households were unaware of the existence of a public pharmacy in their county. Second, even when households knew about the pharmacy, they were not perfectly informed about the lower prices and other attributes. The existence of imperfect information provides us with a unique opportunity to randomly expose consumers to public pharmacies using our experiment, and thus to measure individual responses to them.

The treatment consisted of an informational flyer, displayed in Figure A.10. It provided information about the presence of a public pharmacy in the county and stated that it offered lower prices but longer waiting times than private pharmacies. Also, it included the pharmacy's location, contact information, opening hours, and eligibility requirements. We delivered the flyer to consumers exiting private pharmacies in the 20 counties with public pharmacies in Santiago, displayed in Figure A.11. The information was tailored to each county.

In terms of recruitment, enumerators approached consumers leaving a private pharmacy in each county and assessed their eligibility. Eligible participants were those who (i) lived and were registered to vote in the county, (ii) had purchased a prescription drug, and (iii) were not registered with

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contribute to the low switching rate to public pharmacies, other differences between public and private pharmacies play a relevant role as well. Calculations are available from the authors.

<sup>35</sup>Other unmeasured welfare effects include potential decreases in incentives for R&D. However, we believe that this effect is likely small given the Chilean market represents only a small share of the revenues of the pharmaceutical companies doing R&D.



the public pharmacy. To incentivize participation, everyone who responded to the 5-minute survey automatically entered a lottery for a television set. Overall, 1,855 individuals were approached and 826 enrolled in the study. The baseline survey collected information on awareness of public pharmacies and their attributes, intention to vote for the incumbent mayor in the upcoming election, age, education, and access to the internet, among others. When the survey was completed, participants were randomly assigned to treatment and control groups. The enumerator only learned the assignment of the individual after completing the survey. We conducted this survey between October 12 and 20, 2016, right before the local elections. Figure A.12 summarizes the timeline of the events in the experiment.

Two months after the baseline survey, we conducted a follow-up survey to measure the same variables as in the baseline. We also collected information about their relationship with the public pharmacy in their county. We conducted this survey by phone and were able to complete the survey for 514 participants—almost two-thirds of the sample.<sup>36,37</sup>

Table A.9 compares both groups at baseline. Participants are on average 45 years old and 61 percent of them are female. More than 60 percent work, and most use the internet frequently. Half of the participants planned to vote for the incumbent and almost three out of four reported having participated in the previous election. Slightly less than 70 percent knew about the existence of a public pharmacy. As expected, column (4) shows that almost all variables are balanced across groups. The exception is awareness of the public pharmacy, which we control for in the analysis.

## 5.2 *Experimental Results*

Table 5 presents results from estimating equation (6) for political outcomes. Columns (1) and (4) study self-reported voting behavior. As many as 28 and 26 percent of the control group individuals reported voting for the incumbent mayor and incumbent party, respectively. The reported vote increases by approximately 6 percentage points for the treatment group in both cases. While these point estimates are large in magnitude, they are not statistically significant at conventional levels, with  $p$ -values of 0.21 and 0.12. To increase the precision of the analysis, columns (2) and (5) control for the intention to vote for the mayor at baseline along other covariates, and include

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<sup>36</sup>Table A.7-A shows that attrition was higher among younger participants, males, with higher support for the incumbent, less turnout in the last election, and less knowledge of the public pharmacy. While this changes the sample composition and decreases the statistical power of the experiment, it does not necessarily threaten its internal validity. Table A.7-B shows that all variables remain balanced across groups among non-attriters.

<sup>37</sup>The survey also verified the delivery of the treatment. Table A.8 shows that treated individuals acknowledged receiving information more often than those in the control group, and recalled public pharmacies' being the core of the information content almost twice as often as the latter.

county fixed effects. Treatment effects using this specification remain similar in magnitude but are indeed more precise, with  $p$ -values of 0.06 and 0.11.<sup>38</sup>

Effects on voting behavior are concentrated among individuals from households with members with chronic conditions. Columns (3) and (6) examine these patterns of heterogeneity. Households with someone with a chronic condition report having voted 8 percentage points more for the incumbent, larger than the 2-7 percentage points higher vote share among treated households without a chronic condition. Although the small sample prevents us from rejecting the null of a similar impact across these groups, the result is consistent with the hypothesis that people most affected by the policy are more likely to support the incumbent.

Finally, columns (7)-(9) repeat the previous estimations but now use as dependent variable an indicator that takes the value of one if the person voted in the election. Estimates reveal a positive impact on the probability of turning out to vote—with point estimates similar in magnitude to previous estimates—although in this case none is statistically significant at conventional levels. All in all, these results suggest that awareness of public pharmacies and their characteristics increased consumer support for the incumbent mayor.

We combine these results with estimates of consumer savings from Section 4.3 to estimate the political returns of public pharmacies. The experiment suggests that introducing a public pharmacy increases the number of votes for the incumbent by 1,055, relative to an average of 16,105 total votes across counties in the 2012 local election. Our estimates of the effects on drug expenditure imply that the incumbent obtains 1 additional vote per \$169 of yearly consumer savings. We also consider the monthly savings of consumers who switch to public pharmacies and focus on consumers with chronic conditions. Within that population, the average individual realizes monthly savings of \$45.9. These “transfers” increased political support of the incumbent mayor by 8.1 percentage points. For reference, Manacorda et al. (2011) find that in Uruguay, a targeted monthly transfer of \$70 increased political support for the incumbent government by 11 percentage points.

## 6 Conclusion

State-owned firms compete with the private sector in a variety of markets. The costs and benefits of such competition have been difficult to evaluate empirically. In this paper, we leverage the decentralized entry of state-owned firms to a fully deregulated private market of pharmaceutical retailers. We show that the public option emerged as a low-price and low-quality option and affected the shopping behavior of local consumers, which generated market segmentation and higher

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<sup>38</sup>To account for the effects of attrition, Table 5 presents Lee bounds. The lower bound is positive but not statistically significant and the upper bound is positive and statistically significant across the three outcomes we study.

prices in the private sector. Although public pharmacies created winners and losers within local markets, overall consumer savings outweighed the costs of public pharmacies.

While our study focuses on a particular form of public-private competition, it provides general lessons. First, the equilibrium effects of the public option are shaped by the nature of demand responses. In our context, the public option is less attractive to consumers with a high willingness to pay for service quality relative to drug prices. Market segmentation makes these consumers worse off due to price increases in the private sector.<sup>39</sup> Second, our analysis highlights the fact that public competition may be effective in reducing consumer expenditure. In industries with substantial market power in input and retail markets, retail prices are set at markups over marginal costs. Whenever state-owned firms have higher bargaining power in the input market or decide not to exercise market power in the retail market, they may be able to effectively reduce consumer expenditure. Our setting indeed features these two conditions.

The political rewards of state-owned firms could be interpreted as showing that, as a whole, state-owned firms increased welfare. However, we highlight the fact that recent research shows that people may overvalue policies when they do not internalize the general equilibrium effects that affect them (Dal Bó et al., 2018). Our findings are somewhat consistent with this interpretation since the majority of consumers in the market are worse off after the entry of public pharmacies due to increased private pharmacy prices.<sup>40</sup> These findings demonstrate the need to evaluate the market effect of policies instead of drawing conclusions about their desirability based on voting behavior.

Our analysis leaves many questions for future research. Of particular relevance is understanding the choice of quality among state-owned firms. If the quality of state-owned firms were higher, we would expect more consumers to switch to them and strengthen the equilibrium effects toward the private sector. However, changes in the quality of state-owned firms could influence their targeting properties by modifying the population that adopts them (Kleven and Kopczuk, 2011). Furthermore, it is also possible that a higher quality of state-owned firms triggers other strategic responses in the private sector. In the context of retail, these could include changes in the location, prices, or quality of private stores. Our findings thus call for attention to how the interplay between public and private firm attributes may shape equilibrium effects in the market and determine the overall and distributional impacts of state-owned firms.

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<sup>39</sup>Selection markets, like the market for health insurance, are another important context where the nature of demand responses is key for understanding the general equilibrium effects of the public option. A key feature of those settings would be whether the public option is differentially attractive to consumers with different levels of risk.

<sup>40</sup>Recent work by Illanes and Moshary (2020) on the deregulation of retail liquor markets in Washington state also finds evidence consistent with this phenomenon.

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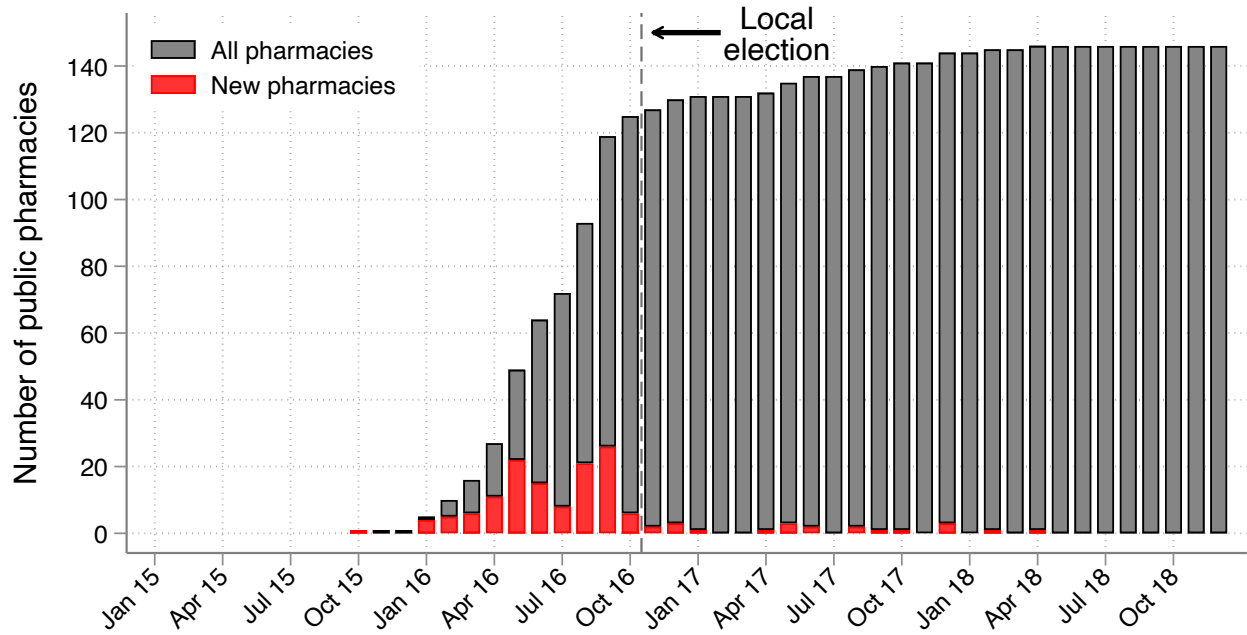
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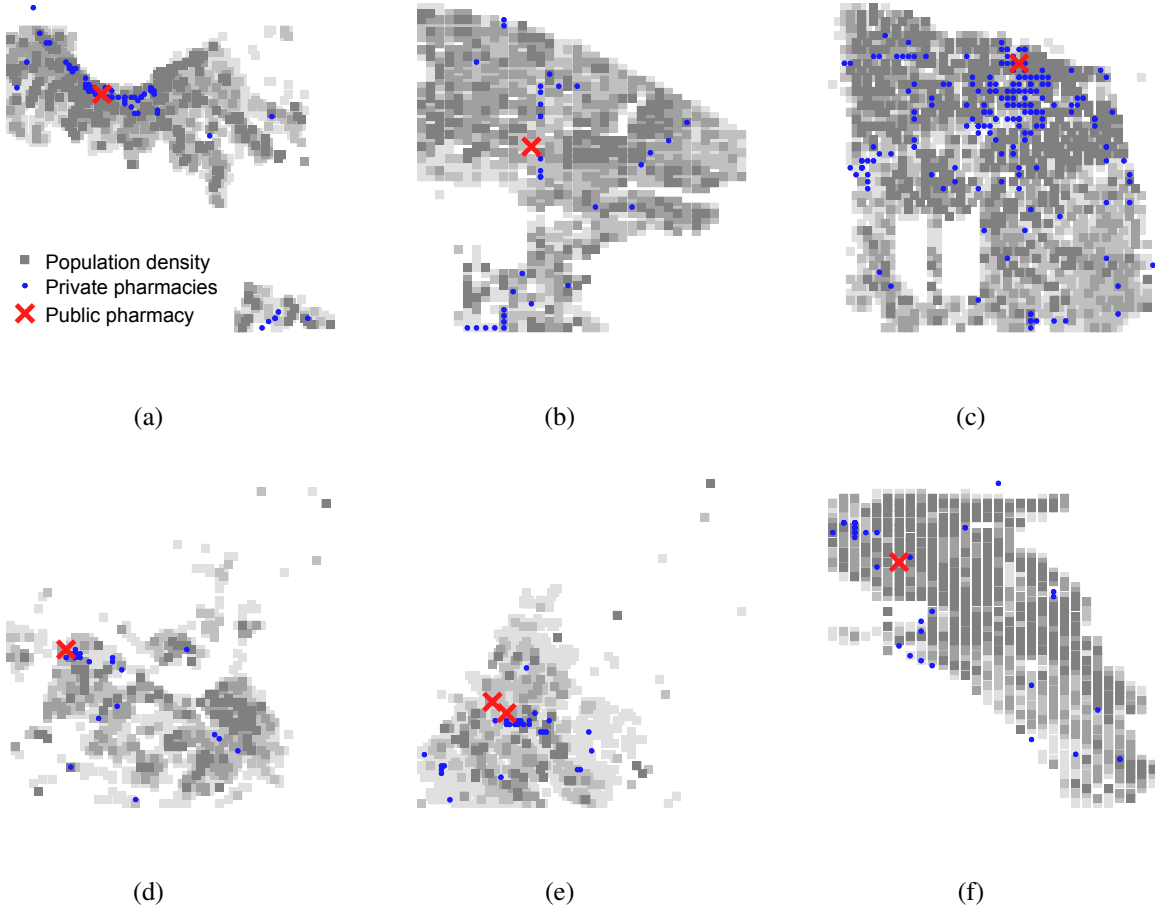
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FIGURE 1: TIMING OF ENTRY OF PUBLIC PHARMACIES



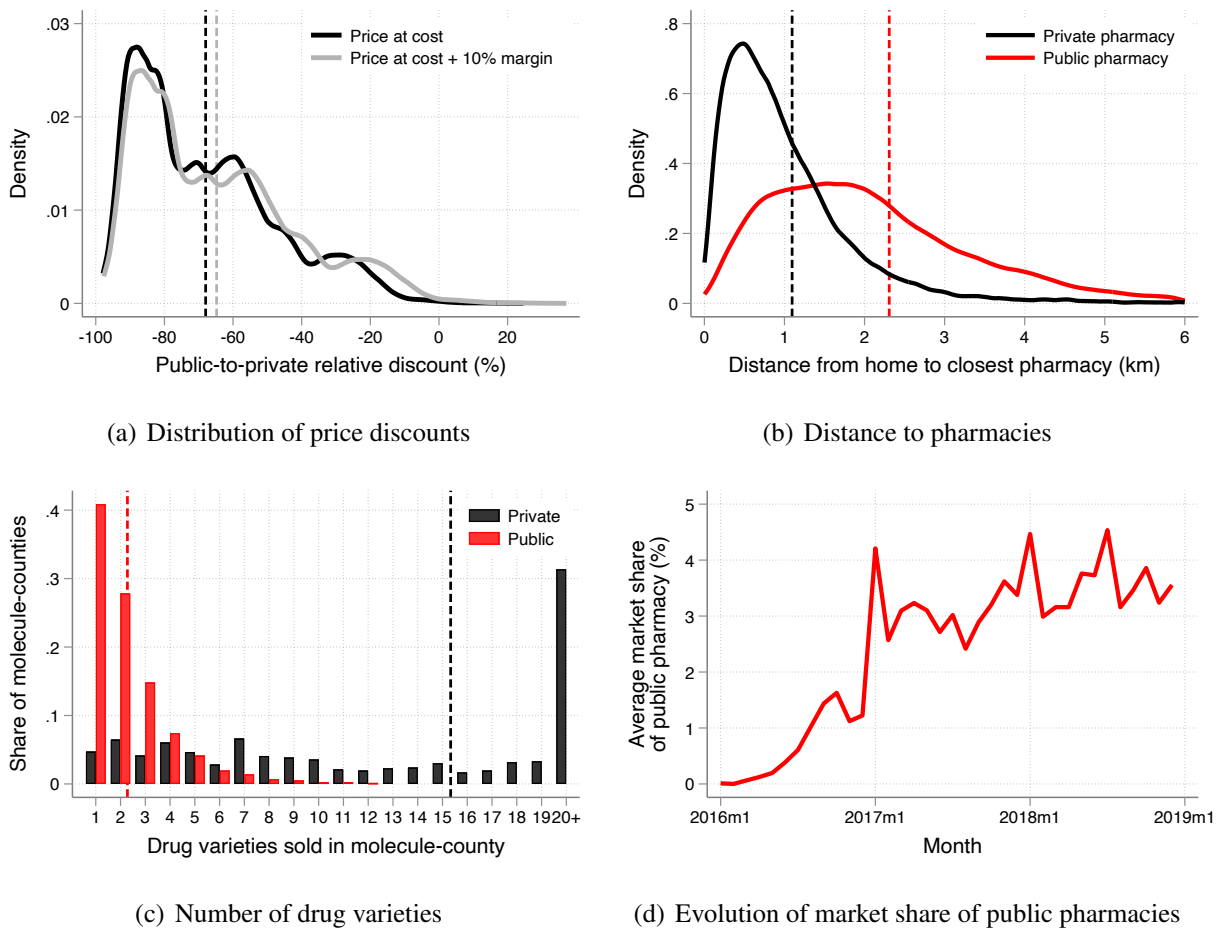
*Notes:* This figure shows the opening dates of public pharmacies (red bars) and the total number of public pharmacies operating (gray bars) in each month between January 2015 and December 2018. The y-axis indicates the total number of public pharmacies opened or the total number of public pharmacies operating each month during this period. The first public pharmacy opened in October 2015. The vertical dashed line in October 2016 indicates the month of the 2016 local election in which most mayors who opened public pharmacies ran for reelection.

FIGURE 2: LOCATIONS OF PUBLIC PHARMACIES IN LOCAL MARKETS



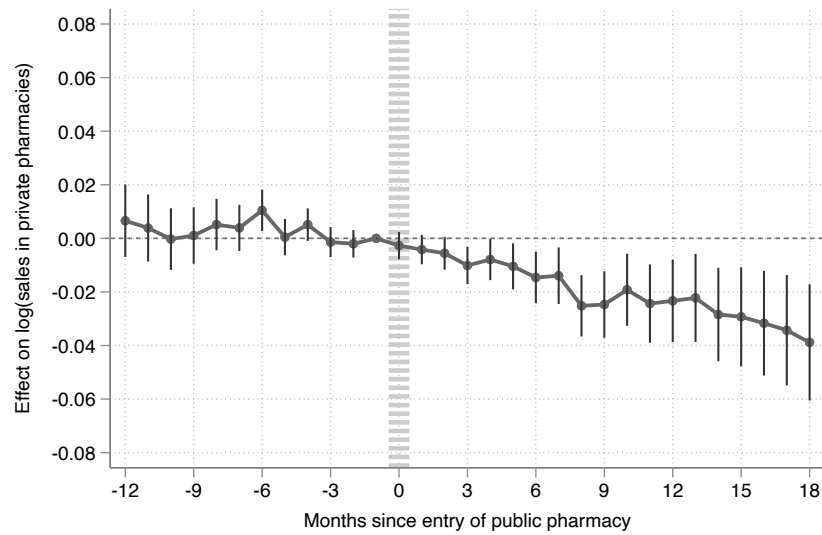
*Notes:* Each map represents a local market defined as a county. The maps display the exact locations of private pharmacies (blue dots), public pharmacies (red cross), and population density in cells of  $111 \times 111$  meters (gray scale). White cells correspond to unpopulated (e.g., parks) or commercial areas. We categorize population density in the following five bins:  $[0, 10)$ ,  $[10, 50)$ ,  $[50, 100)$ ,  $[100, 150)$ , and more than 150 individuals. We use the home addresses of all individuals in the country as revealed by the official Electoral Registry of 2017. The maps correspond to counties in the north, center, and south of the country: (a) Valparaíso, (b) Recoleta, (c) Santiago, (d) Valdivia, (e) Talca, and (f) Iquique.

FIGURE 3: RELATIVE PRICES AND ATTRIBUTES BETWEEN PRIVATE AND PUBLIC PHARMACIES

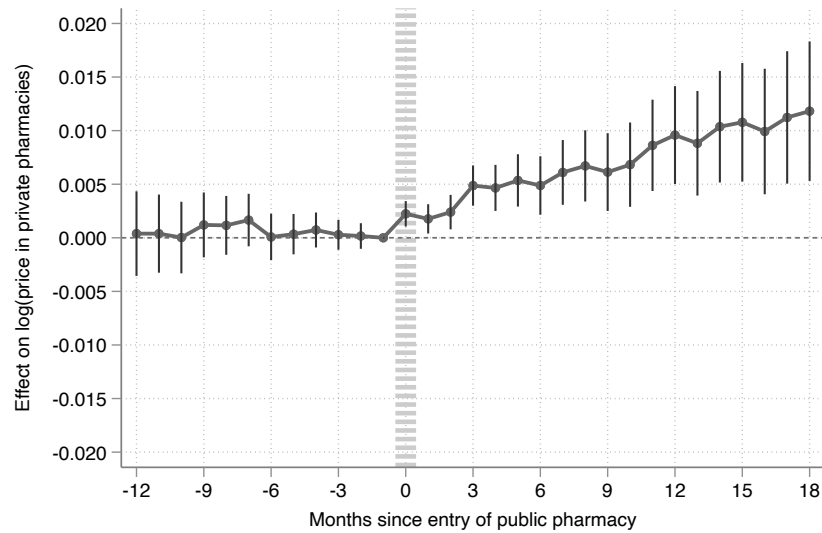


*Notes:* Panel (a) displays the distribution of proportional discounts of drugs at public pharmacies relative to private pharmacies. The plot is computed using a matched sample of the exact same drug observed in both the CENABAST (public pharmacies) and IQVIA (private pharmacies) datasets for a given county and month during 2017–2018. Because the CENABAST data only provide the cost to public pharmacies, we compute price discounts for public pharmacies pricing at cost (black) and at a margin of 10 percent over cost (gray). The dashed vertical lines indicate the mean price discount for each scenario. Panel (b) shows the density of distance from people’s homes to the closest private pharmacy (black) and to the public pharmacy in counties with a public pharmacy. The dashed vertical lines indicate the respective means of both distributions. Panel (c) describes the number of drug presentations of a given molecule sold in a county over 2017–2018 for private (black) and public (red) pharmacies, whenever both private and public pharmacies sell at least one drug of the molecule. Panel (d) displays the average market share of public pharmacies across molecules and counties in each month during 2016–2018.

FIGURE 4: IMPACT OF PUBLIC PHARMACIES ON SALES AND PRICES IN PRIVATE PHARMACIES



(a) Sales



(b) Prices

*Notes:* These figures present event-study estimates of the impact of public pharmacies on private pharmacy sales in Panel (a), and on private pharmacy prices in Panel (b). The unit of observation is a molecule per market in a given month and we use 681,120 observations in panel (a) and 648,885 in panel (b). The empirical strategy uses panel data for the period between 2014 and 2018 and exploits the staggered entry of public pharmacies from October 2015 onward in an event-study design. In Panel (a) the dependent variable is logged sales and in Panel (b) the dependent variable is logged prices. The  $x$ -axis indicates the month with respect to the opening of the public pharmacy, i.e., 18 means 18 months after the opening, and  $-12$  means 12 months before the opening. Dots indicate estimated coefficients and vertical lines indicate the corresponding 95 percent confidence interval.

TABLE 1: DESCRIPTIVE STATISTICS IN COUNTIES WITH AND WITHOUT PUBLIC PHARMACIES

	(1)	(2)	(3)	(4)
	County has public pharmacy			
	Yes	No	Difference (1)–(2)	Timing of entry
Panel A: Pharmacies and hospitals				
Private pharmacies per 100,000 inhabitants	13.59	7.72	5.86***	-0.003
Log sales in private pharmacies	15.37	15.15	0.21**	-0.465
Price index in private pharmacies	931	873	59**	0.001
Hospitalizations per 100,000 inhabitants	9,440	8,126	1,313***	0.00
Deaths per 100,000 inhabitants	209	177	32***	-0.02
Panel B: Socioeconomic characteristics				
Log household income	12.97	12.61	0.36***	-0.467
Age of inhabitants	44.50	45.67	-1.18***	0.115
Average unemployment rate	0.10	0.09	0.01***	7.091
Share with public health insurance	0.83	0.89	-0.06***	1.400
Self reported health (1-7)	5.54	5.49	0.05*	1.900*
Number of doctor visits	0.32	0.30	0.02	1.359
Population (in 10,000)	9.70	1.88	7.82***	-0.425**
Panel C: Political characteristics				
Number of competitors	3.57	3.20	0.37***	0.121
Winning margin	0.19	0.17	0.02	-3.768
Vote share winner	0.54	0.53	0.01	5.951
Incumbent coalition wins	0.62	0.57	0.05	0.439
Incumbent coalition: independent	0.32	0.34	-0.03	-0.045
Incumbent coalition: left-wing	0.47	0.36	0.10*	-1.161**
Incumbent coalition: right-wing	0.22	0.29	-0.07	–
Number of counties	146	198	–	146

*Notes:* This table presents descriptive statistics for counties with and without a public pharmacy in 2018 in columns (1) and (2), respectively. Characteristics in panel A are own construction using data from the Public Health Institute (ISP, DEIS) and IQVIA in 2014. Socioeconomic characteristics in Panel B are own construction using data from the survey National Socioeconomic Characterization (CASEN) conducted in 2015, with the exception of “Population” data, which are publicly available on the website of the National Statistics Bureau (INE). Political characteristics in panel C are own construction using data from the Electoral Service (SERVEL). Column (3) reports the difference between columns (1) and (2) and its statistical significance. Column (4) uses the cross-section of 146 counties with public pharmacies and reports coefficients from an ordered logit using the order in which public pharmacies opened as the dependent variable—the first pharmacy has a value of 1 and the last the value of 146—and all market and political characteristics as explanatory variables. Significance level in columns (3)–(4): \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

TABLE 2: EFFECTS OF PUBLIC PHARMACIES ON DRUG SALES AND PRICES IN THE PRIVATE MARKET

	(1) log(sales)	(2) log(price)
Panel A: Main estimates		
All sample	-0.038*** (0.007)	0.010*** (0.002)
Panel B: Heterogeneity by chronic condition		
Molecules for chronic conditions ( $\beta_{\text{chronic}}$ )	-0.045*** (0.007)	0.010*** (0.002)
Molecules for non-chronic conditions ( $\beta_{\text{non-chronic}}$ )	-0.028*** (0.008)	0.011*** (0.002)
$p$ -value: $\beta_{\text{chronic}} = \beta_{\text{non-chronic}}$	0.006	0.429
Panel C: Heterogeneity by relative product variety		
High public-private variety ratio ( $\beta_{\text{high variety}}$ )	-0.044*** (0.007)	0.013*** (0.002)
Low public-private variety ratio ( $\beta_{\text{low variety}}$ )	-0.033*** (0.008)	0.007*** (0.002)
$p$ -value: $\beta_{\text{high variety}} = \beta_{\text{low variety}}$	0.020	0.000
Panel D: Heterogeneity by distance to private pharmacy		
Private pharmacies are close to public pharmacy ( $\beta_{\text{close}}$ )	-0.044*** (0.008)	0.007*** (0.002)
Private pharmacies are far from public pharmacy ( $\beta_{\text{far}}$ )	-0.032*** (0.007)	0.013*** (0.002)
$p$ -value: $\beta_{\text{close}} = \beta_{\text{far}}$	0.008	0.000
Observations	681,120	649,885
Molecule-by-month FE	Yes	Yes
Molecule-by-market FE	Yes	Yes

*Notes:* This table presents the 18-month effect of the impact of public pharmacies on private pharmacies' sales and prices. These estimates are calculated as  $\bar{E}_{18} \times (\beta^{\text{jump}} + (18 + 1)\beta^{\text{phase in}})$ , where  $\bar{E}_{18}$  is the average share of population across markets with access to a public pharmacy 18 months after the first pharmacy in the local market was introduced. We estimate the on-impact effect  $\beta^{\text{jump}}$  and the trend break effect  $\beta^{\text{phase in}}$  using an exposure difference-in-differences design that leverages the staggered introduction of public pharmacies in the panel data of molecules observed by market and month in the period 2014-2018. We report estimates of  $\beta^{\text{jump}}$  and  $\beta^{\text{phase in}}$  in Table A.2. In Panel B, exposure to public pharmacies is interacted with an indicator for whether a molecule is targeted toward a chronic condition or not. In Panel C, exposure is interacted with an indicator for whether there is a high ratio of variety of products within molecule in public pharmacies relative to private pharmacies defined as above or below the median of the distribution. In Panel D, exposure is interacted with an indicator for whether private pharmacies are located "near" or "far" from public pharmacies. We use the average number of public pharmacies operating within 400 meters of private pharmacies and split the sample in two using the median of this cross-sectional market-level variable. Standard errors clustered at the molecule-by-market level are displayed in parentheses. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

TABLE 3: EFFECTS OF PUBLIC PHARMACIES ON MUNICIPAL FINANCE

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Sub-categories of health related to public pharmacies		All health services		Non-health services		All services	
	Spending	Revenue	Spending	Revenue	Spending	Revenue	Spending	Revenue
Public pharmacy 18-month effect	0.266*** (0.071)	0.187** (0.082)	0.048*** (0.016)	0.034** (0.016)	-0.048 (0.035)	-0.030 (0.034)	0.015 (0.015)	0.010 (0.015)
Avg. dep. var. in 2014	9.144	6.518	182.60	181.43	513.08	548.73	695.68	730.15
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Counties	320	320	321	321	322	322	322	322
Observations	2,200	2,205	2,240	2,240	2,228	2,227	2,243	2,243

*Notes:* This table presents our estimates for the impact of public pharmacies on municipal finances. We observe a panel of counties every year in the period 2013–2019 and exploit the staggered entry of pharmacies in a parametric event study analysis. The dependent variable is the logarithm of total spending (in U.S. dollars) per capita (2013 population) in odd columns and the logarithm of total revenue per capita in even columns. The 18-month effect is the linear combination of regression coefficients  $\pi^{\text{jump}} + (1.5 + 1) \times \pi^{\text{phase in}}$ . Table A.3 presents full regression results, i.e., estimates of  $\pi^{\text{jump}}$  and  $\pi^{\text{phase in}}$ . We focus on 18-month effects to compare the cost of public pharmacies with their impact on sales and prices in private pharmacies (Panel (a) of Table 2). Standard errors clustered at the county level are displayed in parentheses. Significance level: \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .



TABLE 4: EFFECT ON AVOIDABLE HOSPITALIZATIONS ASSOCIATED WITH CHRONIC DISEASES

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Avoidable hospitalizations per 100,000 inhabitants							
	Number of hospitalizations		Days of hospitalizations		Number of surgeries		Number of deaths	
Public pharmacy 18-month effect	-0.891 (0.788)	-1.023 (0.826)	-5.837 (7.815)	-5.061 (8.527)	0.116 (0.175)	0.075 (0.195)	0.092 (0.084)	0.123 (0.093)
Health insurance	All	Public	All	Public	All	Public	All	Public
Mean of dep. var. in 2014	17.93	19.18	158.1	172.5	1.724	1.907	0.736	0.828
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Month fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Counties	344	344	344	344	344	344	344	344
Observations (county-month-years)	28,320	28,320	28,320	28,320	28,320	28,320	28,320	28,320

*Notes:* This table presents our estimates for the impact of public pharmacies on avoidable health outcomes. The outcomes of interest are the number of hospitalizations (columns 1-2), days of hospitalizations (3-4), number of surgeries (columns 5-6), and number of deaths (columns 7-8). For each outcome, the first column uses the count of the outcome per 100,000 inhabitants in a county regardless of individual health insurance, and the second column restricts that count to individuals with publicly provided insurance (FONASA). We observe a panel of counties every month in the period 2013–2019 and exploit the staggered entry of pharmacies in a parametric event study analysis. The 18-month effect is the linear combination of regression coefficients  $\pi^{\text{jump}} + (18 + 1) \times \pi^{\text{phase in}}$ . Table A.5 presents full regression results, i.e., estimates of  $\pi^{\text{jump}}$  and  $\pi^{\text{phase in}}$ . We focus on 18-month effects to use the same horizon of effects as in the previous estimates in the paper. We report the mean of the dependent variable for 2014 among counties that ever introduce a public pharmacy, the year before most public pharmacies entered the market. Standard errors clustered at the county level are displayed in parentheses. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1.

TABLE 5: EXPERIMENTAL RESULTS FOR POLITICAL OUTCOMES

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Voted incumbent mayor			Voted incumbent party			Voted in the election		
Treatment	0.057 (0.045)	0.075* (0.039)		0.064 (0.040)	0.056 (0.035)		0.066 (0.046)	0.052 (0.044)	
Treatment $\times$ chronic ( $\beta_C$ )			0.080 (0.051)			0.081* (0.044)			0.040 (0.055)
Treatment $\times$ non-chronic ( $\beta_{NC}$ )			0.067 (0.065)			0.020 (0.058)			0.068 (0.073)
Dependent variable at baseline		0.366*** (0.051)	0.367*** (0.051)		0.348*** (0.048)	0.350*** (0.048)		0.418*** (0.052)	0.416*** (0.052)
Lee bounds		[0.033, 0.182***]			[0.048, 0.170***]			[0.014, 0.159**]	
$p$ -value for $H_0: \beta_C = \beta_{NC}$	-	-	0.883	-	-	0.408	-	-	0.763
Mean for control group	0.281	0.277	0.277	0.263	0.255	0.255	0.541	0.524	0.524
Observations	398	368	368	475	435	435	475	435	435
R-squared	0.004	0.515	0.515	0.005	0.488	0.488	0.004	0.641	0.641
Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
County FE	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes

*Notes:* This table presents our estimates of the political impact of public pharmacies using data from the field experiment described in Section 5. The unit of observation is an individual who buys pharmaceuticals at private pharmacies in the capital city of Santiago. The treatment is information about public pharmacies delivered in the form of a flyer by enumerators after completing the baseline survey in October 2016, before the local election. All dependent variables were measured in follow-up surveys conducted in December 2016, after the local election. We present cross-sectional results using three specifications, one without controls (columns 1, 4, and 7), one with controls (columns 2, 5, and 8), and one with controls and interacting the treatment with an indicator for individuals with a chronic condition (columns 3, 6, and 9). The set of control variables includes age and indicators for chronic condition, having completed high school education, female, and public insurance. Reported Lee bounds are computed using only the treatment indicator as covariate. Robust standard errors in parentheses. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

## ONLINE APPENDIX

### *The Economics of the Public Option: Evidence from Local Pharmaceutical Markets*

Juan Pablo Atal, José Ignacio Cuesta, Felipe González, and Cristóbal Otero.

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## Online Appendix A The determinants of retail drug prices

In the main text, we argue that two conditions that generate price differences between state-owned and private firms are the higher bargaining power of the former in the wholesale market and the exercise of market power of the latter in the retail market. In this section, we present a model that formalizes this intuition.

### A.1 Setup

We consider a sequential monopoly model with Nash bargaining. An upstream monopoly produces a drug that is sold to a retail pharmacy that is a downstream monopoly. The model allows for this downstream firm to represent the private pharmacy, the public pharmacy, or some combination between them—we specify how the downstream firm’s objective function captures these possibilities below. The marginal cost of the upstream monopoly is  $c$  and the wholesale price the retailer pays is  $t$ . There are no additional marginal costs downstream.

We start by introducing the objective functions of the upstream firm and the retailer. The upstream monopoly maximizes profits:

$$\Pi_U(t) = (t - c)\bar{q}(t),$$

where  $\bar{q}(t) \equiv q(p(t))$  are the sales that result when the downstream retailer chooses the optimal retail price given the wholesale price  $t$ .

The downstream firm sets prices by taking into account both profits and consumer surplus, with a weight on consumer surplus equal to  $\lambda$ . Omitting the dependence of prices with respect to the wholesale price, the objective function of the retailer is:

$$V_D(p) = (p - t)q(p) + \lambda CS(p),$$

where  $q(p)$  is the demand function, for which we assume  $q'(p) < 0$  and  $q''(p) \geq 0$ . The parameter  $\lambda$  measures the degree of alignment between the retailer and consumers. If  $\lambda < 1$ , the retailer values profits more than consumer welfare;  $\lambda > 1$  implies that the retailer values consumer welfare more than profits; and  $\lambda = 1$  means that consumer welfare and profits are valued equally by the retailer and hence that the retailer maximizes total welfare.<sup>41</sup> In terms of the downstream market structure, this specification of the retailer objective is akin to a mixed oligopoly model for the retail market in which private and state-owned firms compete (see, e.g., Merrill and Schneider 1966; Beato and Mas-Colell 1984; De Fraja and Delbono 1989; Cremer et al. 1991; Duarte et al. 2021).

**Bargaining over wholesale price.** The upstream and downstream firms bargain over wholesale prices. The wholesale price  $t$  maximizes the Nash product of the gains from trade for both firms:

$$V_D(p(t))^\zeta \times (\Pi_U(t))^{1-\zeta},$$

---

<sup>41</sup>See also Timmins (2002) and Gowrisankaran et al. (2015) for similar specifications of firm objectives when aligned with consumers.

where  $\zeta$  is the bargaining power of the retailer.

**Optimal pricing upstream and downstream.** The first-order condition of the Nash bargaining problem is:

$$(t - c)q'(p)p'(t) + q = \left( \frac{\zeta}{1 - \zeta} \right) \frac{t - c}{(p - t) + \lambda \frac{CS}{q}} \times q, \quad (4)$$

where it is useful to note that this equation simplifies to the standard first order condition of the bilateral monopoly model in the case of  $\lambda = 0$ , where the retailer places no weight on consumer surplus (Lee et al., 2021).

The optimal retailer price is given by:

$$p = t - \frac{q}{q'} - \lambda \frac{CS'}{q'},$$

which, by using the fact that  $CS' = -q(p)$ , simplifies to:

$$p = t - (1 - \lambda) \frac{q}{q'},$$

which only holds when  $\lambda < 1$ . When  $\lambda \geq 1$ , the downstream firm is at a corner solution where it sets prices at marginal cost, namely  $p = t$ . Overall, the optimal price downstream is given by:

$$p = \begin{cases} t - (1 - \lambda) \frac{q}{q'} & \lambda < 1 \\ t & \lambda \geq 1. \end{cases} \quad (5)$$

Market outcomes are jointly determined by equations (4) and (5), and depend on the bargaining power of the retailer and the extent to which the retailer is aligned with consumers and value consumer surplus.

## A.2 Comparative Statics

In this section, we deliver the main results of the model. In particular, we show how wholesale and retail prices vary with the retailer's bargaining power and market power, which depend on the parameters  $\zeta$  and  $\lambda$ , respectively. These are the results that map to the two conditions we discuss in the main text for why public state-owned firms may offer lower prices than private firms in our setting. We start by introducing three assumptions:

**Assumption 1** (Decreasing Marginal Revenue). *Marginal revenue  $MR(q) = p(q) + qp'(q)$  is decreasing in  $q$ , where  $p(q)$  is the inverse demand curve.*

**Assumption 2.**  $\frac{qq''}{q'^2}$  is weakly increasing in  $p$ .

**Assumption 3.**  $\frac{q^2}{-q'} - CS \geq 0$ .

These assumptions provide conditions under which the two comparative statics of interest hold. Assumption 1 guarantees the existence of a profit-maximizing price for a monopolist facing a convex cost function and is implied by log-concavity of demand (see e.g., Kang and Vasserman 2022). Assumption 3 is also implied by log-concavity, as shown in Section A.4.1. Log concavity is a commonly-used assumption in industrial organization, and hence it is not particularly restrictive (Bagnoli and Bergstrom, 2006). This property of demand ensures that the first order condition of the monopoly is sufficient for profit maximization.

We start by establishing general results for how market outcomes vary with the degree of bargaining power downstream,  $\zeta$ . Lemma 1 shows that under Assumption 1 and Assumption 2, wholesale prices and downstream prices are decreasing on the retailer's bargaining power  $\zeta$ .

**Lemma 1.** *Wholesale prices and retail prices are decreasing in the bargaining power of the retailer. For  $\lambda \geq 1$  and if Assumption 1 holds, then  $\partial t / \partial \zeta < 0$  and  $\partial p / \partial \zeta < 0$ . For  $\lambda < 1$  and if Assumption 1 and 2 hold, then  $\partial t / \partial \zeta < 0$  and  $\partial p / \partial \zeta < 0$ .*

*Proof.* See Section A.4.3 □

We now establish general results for how market outcomes vary with the extent of alignment between the retailer and consumers,  $\lambda$ . When  $\lambda \geq 1$ , the retailer sets its price to be equal to the wholesale price,  $p = t$ . Lemma 2 shows that in this case, the wholesale price and the retail price are independent of  $\lambda$ . When  $\lambda < 1$ , the wholesale price is not always decreasing with  $\lambda$ . The intuition is as follows: as  $\lambda$  goes up, the retailer would like to give away profits to increase output. In some cases, this allows the upstream firm to set a higher wholesale price. Regardless, Lemma 2 shows that retail prices are decreasing with  $\lambda$  under Assumptions 1, 2 and 3, which is the result of main interest in our context.

**Lemma 2.** *The retail price is weakly decreasing in the weight given to consumer surplus,  $\lambda$ . In particular, for  $\lambda \geq 1$  we show that  $\partial p / \partial \lambda = 0$  and  $\partial t / \partial \lambda = 0$ . For  $\lambda < 1$  and if Assumptions 1, 2, and 3 hold, then  $\frac{\partial p}{\partial \lambda} < 0$ .*

*Proof.* See Section A.4.4. □

### A.3 Parametric Examples

Lemmas 1 and 2 provide general conditions under which retail prices are lower when retailers have more bargaining power, and when retailers are more aligned with consumers. These conditions hold for multiple families of demand that satisfy combinations of Assumptions 1, 2, and 3. To provide examples for these results, Lemmas 3-7 show that retail prices are weakly decreasing with  $\zeta$  and  $\lambda$  for commonly used families of demand functions.

**Lemma 3** (CES demand). *Consider the CES demand function of the form  $q = p^\alpha$ , with  $\alpha < -1$ . With CES demand, wholesale and retail prices are weakly decreasing in the bargaining power downstream and in the weight given to consumer surplus. For  $\lambda < 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} < 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ . For  $\lambda \geq 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} = 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ .*

*Proof.* See Section A.4.5. □

**Lemma 4** (Constant marginal revenue). *Consider a demand function that features a constant marginal revenue curve  $q = \frac{1}{p-a}$  (CMR demand). With CMR demand, wholesale prices and retail prices are weakly decreasing in the bargaining power downstream and in the weight given to consumer surplus. For  $\lambda < 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} < 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} < 0$ . For  $\lambda \geq 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} = 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ .*

*Proof.* See Section A.4.6. □

**Lemma 5** (Logit demand). *Consider a logit demand function  $q = \frac{e^{-\beta p}}{1+e^{-\beta p}}$ . With logit demand, retailer prices are weakly decreasing in the retailer's bargaining power and in the weight given to consumer surplus. For  $\lambda < 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} < 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$ . For  $\lambda \geq 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} = 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ .*

*Proof.* See Section A.4.7. □

**Lemma 6** (Exponential demand). *Consider an exponential demand function  $q = e^{-\beta p}$ . With exponential demand, retail prices are weakly decreasing in the retailer's bargaining power and in the weight given to consumer surplus. For  $\lambda < 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} < 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$ . For  $\lambda \geq 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} = 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ .*

*Proof.* See Section A.4.8. □

**Lemma 7** ( $\rho$ -linear demand). *Consider a  $\rho$ -linear demand function  $q = (a - bp)^{1/\rho}$ . With  $\rho$ -linear demand, retail prices are weakly decreasing in the retailer's bargaining power and in the weight given to consumer surplus. For  $\lambda < 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} < 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$ . For  $\lambda \geq 1$ ,  $\frac{\partial p}{\partial \zeta} < 0$  and  $\frac{\partial p}{\partial \lambda} = 0$ , and in addition  $\frac{\partial t}{\partial \zeta} < 0$  and  $\frac{\partial t}{\partial \lambda} = 0$ .*

*Proof.* See Section A.4.9. □

## A.4 Additional Lemmas and Proofs

### A.4.1 Assumption 3 and log-concavity

**Lemma 8.** *If  $q$  is twice differentiable and log-concave, then Assumption 3 holds:  $\frac{q^2}{-q'} - CS \geq 0$ .*

*Proof.* Since  $q$  is differentiable,  $q'$  exists and is finite.  $\frac{q^2}{-q'} = 0$  and  $CS = 0$  if  $q = 0$ . As  $\lim_{p \rightarrow +\infty} q = 0$ ,  $\lim_{p \rightarrow \infty} \frac{q^2}{-q'} - CS = 0$ . Taking the derivatives of  $f(p) := \frac{q^2}{-q'} - CS$ , we get:

$$f'(p) = \frac{-2qq'^2 + q^2q''}{q'^2} + q = \frac{-qq'^2 + q^2q''}{q'^2} = q \frac{-q'^2 + qq''}{q'^2} < 0,$$

as  $q$  is log-concave. So  $f(p)$  is decreasing in  $p$ . From  $\lim_{p \rightarrow \infty} f(p) = 0$  we get  $f(p) \geq 0$ . □

#### A.4.2 Decreasing Marginal Revenue

We provide an equivalent expression of decreasing marginal revenue for a twice-differentiable function.

**Lemma 9.** *If  $q$  is twice differentiable, then  $2q'^2 - qq'' \geq 0$  if and only if  $q$  has decreasing marginal revenue.*

*Proof.* Rewrite marginal revenue  $MR$  as a function of  $p$  by inverse function theorem:

$$MR(p) = p + \frac{q(p)}{q'(p)}.$$

Taking the derivative with respect to  $p$  yields:

$$MR'(p) = 1 + \frac{q'^2 - qq''}{q'^2} = \frac{2q'^2 - qq''}{q'^2}.$$

so that marginal revenue is increasing in  $p$  if and only if  $2q'^2 - qq'' \geq 0$ . Since  $q$  is decreasing in  $p$ , marginal revenue is decreasing in  $q$  if and only if  $2q'^2 - qq'' \geq 0$ .  $\square$

#### A.4.3 Proof of Lemma 1

Case 1:  $\lambda < 1$  In this case, the first order condition for the retailer holds and therefore:

$$F_2 := p - t + (1 - \lambda) \frac{q}{q'} = 0.$$

Taking the derivatives with respect to  $t$  yields:

$$\frac{dp}{dt} = -\frac{\frac{\partial F_2}{\partial t}}{\frac{\partial F_2}{\partial p}} = \frac{1}{\lambda + (1 - \lambda) \frac{2(q')^2 - qq''}{(q')^2}}.$$

By Assumption 1,  $\frac{dp}{dt} > 0$ . Imposing condition  $F_2$  on Equation (4), we obtain:

$$F_1 := -\frac{q}{\frac{q'p'}{q} + \frac{1}{t-c}} + \frac{1-\zeta}{\zeta} \left[ -(1-\lambda) \frac{q^2}{q'} + \lambda CS \right] = 0,$$



such that:

$$\begin{aligned}\frac{\partial F_1}{\partial \zeta} &= -\frac{1}{\zeta^2} \left[ -(1-\lambda) \frac{q^2}{q'} + \lambda CS \right] \\ \frac{\partial F_1}{\partial t} &= -\frac{\frac{(2q'^2 - qq'')p'^2 - qq'p''}{q} + \frac{q'p'(t-c)+q}{(t-c)^2}}{\left( \frac{q'p'}{q} + \frac{1}{t-c} \right)^2} + \frac{1-\zeta}{\zeta} \left[ -(1-\lambda) \frac{qp'(2q'^2 - qq'')}{q'^2} - \lambda qp' \right].\end{aligned}$$

It follows immediately that  $\frac{\partial F_1}{\partial \zeta} < 0$ . The sign of  $p''$  is determined by  $\frac{d \frac{qq''}{q'^2}}{dp}$  since:

$$p'' = \frac{(1-\lambda) \frac{d \frac{qq''}{q'^2}}{dp} \frac{dp}{d\lambda}}{[(2-\lambda) - (1-\lambda) \frac{qq''}{q'^2}]^2}.$$

From Assumption 2,  $\frac{d \frac{qq''}{q'^2}}{dp} \geq 0$ , and therefore  $p'' \geq 0$ . The first term of  $\frac{\partial F_1}{\partial \zeta}$  is weakly negative, and the second term is negative, so  $\frac{\partial F_1}{\partial \zeta} < 0$ . Therefore  $\frac{\partial t}{\partial \zeta} = -\frac{\frac{\partial F_1}{\partial \zeta}}{\frac{\partial F_1}{\partial t}} < 0$ . In addition, we get  $\frac{\partial p}{\partial \zeta} < 0$  since  $\frac{dp}{dt} > 0$ .

Case 2:  $\lambda \geq 1$  With a sufficiently high weight given to consumer surplus, in particular when  $\lambda > 1$ , the retailer will set the price equal to its marginal cost, as shown by equation (5). The Nash bargaining first-order condition in equation (4) becomes:

$$F := \frac{-q^2(t)(t-c)}{(t-c)q'(t) + q(t)} + \frac{1-\zeta}{\zeta} CS(t) = 0.$$

Taking the partial derivative with respect to  $\zeta$  yields:

$$\frac{\partial F}{\partial \zeta} = -\frac{1}{\zeta^2} CS(t) < 0; \quad \frac{\partial F}{\partial t} = -\frac{\frac{2q'^2 - qq''}{q} + \frac{q'(t-c)+q}{(t-c)^2}}{\left( \frac{q'}{q} + \frac{1}{t-c} \right)^2} - \frac{1-\zeta}{\zeta} q < 0.$$

and it follows that under Assumption 1 that  $\partial t / \partial \zeta < 0$ .

#### A.4.4 Proof of Lemma 2

Case 1:  $\lambda \geq 1$  For  $\lambda \geq 1$ , the retailer sets price equal to marginal cost,  $p = t$ . The Nash bargaining first order condition is:

$$F := \frac{-q^2(t)(t-c)}{(t-c)q'(t) + q(t)} + \frac{1-\zeta}{\zeta} CS(t) = 0,$$

which does not contain  $\lambda$ , so that  $t$  does not depend on  $\lambda$ . Thus  $\frac{\partial t}{\partial \lambda} = 0$ . From  $p = t$ , we have  $\frac{\partial p}{\partial \lambda} = 0$ .

Case 2:  $\lambda < 1; \zeta = 0$  In the special case in which  $\zeta = 0$ , the upstream firm acts as a monopoly and sets the wholesale price to maximize its profits. In this case, the upstream firm and retailer profit functions become:

$$\begin{aligned}\Pi_U &= (t - c)q(p) \\ V_D &= (p - t)q(p) + \lambda \text{CS}(p),\end{aligned}$$

and the upstream firm and retailer first order conditions become:

$$\begin{aligned}F_1 &:= (t - c)q'p' + q = 0 \\ F_2 &:= p - t + (1 - \lambda)\frac{q}{q'} = 0,\end{aligned}$$

such that from the retailer's first order condition we obtain:

$$p' = -\frac{\frac{\partial F_2}{\partial t}}{\frac{\partial F_2}{\partial p}} = \frac{1}{1 + (1 - \lambda)\frac{q'^2 - qq''}{q'^2}},$$

which we plug into the upstream firm's first order condition to rewrite  $F_1$  as:

$$q + (t - c)q' \frac{1}{1 + (1 - \lambda)\frac{q'^2 - qq''}{q'^2}} = 0.$$

By combining the two first order conditions, we get:

$$F := \begin{bmatrix} q + (t - c)q' \frac{1}{1 + (1 - \lambda)\frac{q'^2 - qq''}{q'^2}} \\ p - t + (1 - \lambda)\frac{q}{q'} \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix}.$$

Note that:

$$\begin{bmatrix} \frac{\partial F_1}{\partial \lambda} \\ \frac{\partial F_2}{\partial \lambda} \end{bmatrix} = \begin{bmatrix} (t - c)q'p'^2 \frac{q'^2 - qq''}{q'^2} \\ -\frac{q}{q'} \end{bmatrix},$$

and the Jacobian matrix of  $F$  is:

$$J := \begin{bmatrix} \frac{\partial F_1}{\partial t} & \frac{\partial F_1}{\partial p} \\ \frac{\partial F_2}{\partial t} & \frac{\partial F_2}{\partial p} \end{bmatrix} \begin{bmatrix} q'p' & q' + (t - c) \left[ q''p' + q'p'^2(1 - \lambda)\frac{d\frac{q''}{q'^2}}{dp} \right] \\ -1 & \frac{1}{p'} \end{bmatrix},$$

while the determinant of  $J$  is:

$$\begin{aligned} \det(J) &= q' + (t - c) \left[ q'' p' + q' p'^2 (1 - \lambda) \frac{d \frac{q q''}{q'^2}}{dp} \right] + \frac{q' p'}{p'} \\ &= \frac{2q'^2 - q q''}{q'} - (1 - \lambda) q p' \frac{d \frac{q q''}{q'^2}}{dp}. \end{aligned}$$

From Assumption 1 and Lemma 9,  $2q'^2 - q q'' \geq 0$ . This yields  $\frac{2q'^2 - q q''}{q'} \leq 0$ . From assumption

$2, \frac{d \frac{q q''}{q'^2}}{dp} \geq 0$ . So  $\det(J) \leq 0$ .

The inverse matrix of  $J$  is:

$$J^{-1} = \frac{1}{\det(J)} \begin{bmatrix} \frac{1}{p'} & -q' - (t - c) \left[ q'' p' + q' p'^2 (1 - \lambda) \frac{d \frac{q q''}{q'^2}}{dp} \right] \\ 1 & q' p' \end{bmatrix},$$

and using the implicit function theorem we show that:

$$\begin{aligned} \frac{\partial p}{\partial \lambda} &= -\frac{1}{\det(J)} \left[ q' p' \cdot \left( -\frac{q}{q'} \right) + (t - c) q' p'^2 \frac{q'^2 - q q''}{q'^2} \right] \\ &= \frac{1}{\det(J)} q p' \frac{2q'^2 - q q''}{q'^2} < 0. \end{aligned}$$

Case 3:  $\lambda < 1; \zeta < 1$  Rewrite the first order condition for the bargaining problem as:

$$\frac{\partial \pi_u}{\partial t} - \frac{\zeta}{1 - \zeta} \frac{q}{V_D} = 0 \implies (1 - \zeta) F_2 - \zeta q \frac{\pi_u}{V_D} = 0.$$

where  $F_2 := q + (t - c) q'(p) p'(t)$ . The first order condition of the retailer is:

$$F_1 := p - t + (1 - \lambda) \frac{q}{q'} = 0.$$

Combining both conditions yields:

$$F := \begin{bmatrix} F_1 \\ (1 - \zeta) F_2 - \zeta q \frac{\pi_u}{V_D} \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix},$$

for which the partial derivative with respect to  $\lambda$  is:

$$\frac{\partial F}{\partial \lambda} = \begin{bmatrix} \frac{\partial F_1}{\partial \lambda} \\ (1 - \zeta) \frac{\partial F_2}{\partial \lambda} - \zeta q \frac{\partial \frac{\pi_u}{V_D}}{\partial \lambda} \end{bmatrix}.$$

and the Jacobian is:

$$J = \begin{bmatrix} \frac{\partial F_1}{\partial t} & \frac{\partial F_1}{\partial p} \\ (1 - \zeta) \frac{\partial F_2}{\partial t} - \zeta q \frac{\partial \pi_u}{\partial t} & (1 - \zeta) \frac{\partial F_2}{\partial p} - \zeta q' \frac{\pi_u}{V_D} - \zeta q \frac{\partial \pi_u}{\partial p} \end{bmatrix},$$

for which the determinant is:

$$\det(J) = (1 - \zeta) \left( \frac{\partial F_1}{\partial t} \frac{\partial F_2}{\partial p} - \frac{\partial F_1}{\partial p} \frac{\partial F_2}{\partial t} \right) + \zeta \left[ q' \frac{\pi_u}{V_D} + q \frac{\partial \pi_u}{\partial p} + q \frac{\partial \pi_u}{\partial t} \frac{1}{p'(t)} \right].$$

We know that when  $\zeta = 0$ , then  $\frac{\partial F_1}{\partial t} \frac{\partial F_2}{\partial p} - \frac{\partial F_1}{\partial p} \frac{\partial F_2}{\partial t} > 0$ . So we focus on  $M := q' \frac{\pi_u}{V_D} + q \frac{\partial \pi_u}{\partial p} + q \frac{\partial \pi_u}{\partial t} \frac{1}{p'(t)}$ , which can be simplified to  $M = \frac{q}{\zeta V_D p'} [(1 + \zeta) q' p'(t - c) + q]$ . This yields:

$$\det(J) = (1 - \zeta) \left( \frac{\partial F_1}{\partial t} \frac{\partial F_2}{\partial p} - \frac{\partial F_1}{\partial p} \frac{\partial F_2}{\partial t} \right) + \frac{q}{V_D p'} [(1 + \zeta) q' p'(t - c) + q].$$

where the first term is greater than 0 given  $\zeta = 0$ . The second term is decreasing in  $\zeta$ . When  $\zeta \rightarrow 1$ ,  $t \rightarrow c$  because the wholesaler's profit has zero weight in the bargaining stage. Thus the second term is equal to  $\frac{q^2}{V_D p'} > 0$ . So  $|J| > 0$  for all  $\zeta$ . Thus, the inverse of the Jacobian is:

$$J^{-1} = \frac{1}{\det(J)} \begin{bmatrix} (1 - \zeta) \frac{\partial F_2}{\partial p} - \zeta q' \frac{\pi_u}{V_D} - \zeta q \frac{\partial \pi_u}{\partial p} & -\frac{\partial F_1}{\partial p} \\ -(1 - \zeta) \frac{\partial F_2}{\partial t} + \zeta q \frac{\partial \pi_u}{\partial t} & \frac{\partial F_1}{\partial t} \end{bmatrix}.$$

Using these results, we can write the partial derivative of retail price with respect to  $\lambda$  as:

$$\frac{\partial p}{\partial \lambda} = -\frac{1}{\det(J)} \left[ (1 - \zeta) \left( -\frac{\partial F_2}{\partial t} \frac{\partial F_1}{\partial \lambda} + \frac{\partial F_1 \partial F_2}{\partial t} \right) + \zeta q \left( \frac{\partial \pi_u}{\partial t} \frac{\partial F_1}{\partial \lambda} - \frac{\partial \pi_u}{\partial \lambda} \frac{\partial F_1}{\partial t} \right) \right],$$

where since  $\zeta = 0$  we know that  $-\frac{\partial F_2}{\partial t} \frac{\partial F_1}{\partial \lambda} + \frac{\partial F_1 \partial F_2}{\partial t} > 0$ , so we can focus on the sign of the last term:

$$\frac{\partial \pi_u}{\partial t} \frac{\partial F_1}{\partial \lambda} - \frac{\partial \pi_u}{\partial \lambda} \frac{\partial F_1}{\partial t} = \left[ -\frac{q}{q'} \frac{q(V_D + \pi_u)}{V_D^2} - \frac{CS \pi_u}{V_D^2} \right] = \frac{1}{V_D^2} \left[ -\frac{q^2}{q'} V_D + \pi_u \left( -\frac{q^2}{q'} - CS \right) \right],$$

and from Assumption 3, we have  $N > 0$ . Therefore,  $\frac{\partial p}{\partial \lambda} < 0$ .

Case 4:  $\lambda < 1; \zeta = 1$  In this case, the upstream firm will set the wholesale price equal to the marginal cost,  $t = c$ . Equation (5) can be written as:

$$F := (p - c)q' + (1 - \lambda)q = 0.$$

from where by taking partial derivatives with respect to  $p$  we get:

$$\begin{aligned}\frac{\partial F}{\partial p} &= [q' + (p - c)q''] + (1 - \lambda)q' = \lambda q' + (1 - \lambda)\frac{2q'^2 - qq''}{q'} < 0 \\ \frac{\partial F}{\partial \lambda} &= -q < 0,\end{aligned}$$

such that under Assumption 1,  $\frac{\partial p}{\partial \lambda} = -\frac{\frac{\partial F}{\partial \lambda}}{\frac{\partial F}{\partial p}} < 0$ .

#### A.4.5 Proof for Lemma 3 (CES demand)

Notice that the CES function is not quasi-concave. Note also that Assumption 1 holds:

$$2q'^2 - qq'' = 2\alpha^2 p^{2\alpha-2} - \alpha(\alpha - 1)p^{2\alpha-2} = \alpha(\alpha + 1)p^{2\alpha-2} > 0,$$

and that Assumption 2 holds:

$$(\log q)' + (\log q'')' - 2(\log(-q'))' = \log(-\alpha) + \log(1 - \alpha) - 2\log(-\alpha) = \log(1 - \alpha) - \log(-\alpha) > 0,$$

such that Lemma 1 implies that  $\partial p/\partial \zeta < 0$  and  $\partial t/\partial \zeta < 0$ . However, Assumption 3 fails to hold since:

$$\frac{p^{2\alpha}}{-\alpha p^{\alpha-1}} + \frac{p^{\alpha+1}}{\alpha + 1} = \frac{p^{\alpha+1}}{-\alpha} + \frac{p^{\alpha+1}}{\alpha + 1} = \frac{-p^{\alpha+1}}{\alpha(\alpha + 1)} < 0.$$

From equation (5) we get:

$$p = \frac{\alpha}{\alpha + (1 - \lambda)}t,$$

and then from equation (4) we get:

$$(t - c)\alpha p^{\alpha-1} \frac{\alpha}{\alpha + (1 - \lambda)} + p^\alpha = \frac{\zeta}{1 - \zeta} \frac{t - c}{\frac{\lambda-1}{\alpha+1-\lambda}t + \lambda(-\frac{p}{\alpha+1})},$$

which can be simplified to:

$$\alpha^2(t - c) + \alpha t + \frac{\zeta}{1 - \zeta} (\alpha^2 + \alpha)(t - c) = 0,$$

from where it follows that  $t$  is independent of  $\lambda$ , and so  $\frac{\partial t}{\partial \lambda} = 0$  and  $\frac{\partial p}{\partial \lambda} = \frac{\alpha}{(\alpha + (1 - \lambda))^2}t < 0$ .

#### A.4.6 Proof for Lemma 4 (CMR demand)

Notice that the CMR demand is not quasi-concave. Note also that Assumption 1 holds:

$$2q'^2 - qq'' = \frac{2}{(p - a)^4} - \frac{2}{(p - a)^4} = 0,$$

and that Assumption 2 also holds given:

$$\log q + \log q'' - 2 \log q' = -\log(p - a) + \log 2 - 3 \log(p - a) + 4 \log(p - a) = \log 2$$

is constant on  $p$ , i.e., weakly increasing in  $p$ . Then from Lemma 1,  $dp/d\zeta < 0$  and  $dt/d\zeta < 0$ . However, Assumption 3 fails to hold since:

$$\frac{q^2}{-q'} - CS = 1 + \log(p - a).$$

We now check the sign of  $\frac{\partial p}{\partial \lambda}$  when  $\lambda \leq 1$ . The first order condition for the Nash problem in equation (4)) implies:

$$F := \frac{\zeta}{1 - \zeta} \left( \frac{t - c}{c - a} \right) - \frac{1 - \lambda}{\lambda} - \log \lambda + \log(t - a) - C = 0,$$

where  $C$  is an arbitrary constant that nonetheless determines the price. Taking partial derivatives yields:

$$\begin{aligned} \frac{\partial F}{\partial t} &= \frac{\zeta}{(1 - \zeta)(c - a)} + \frac{1}{t - a} > 0 \\ \frac{\partial F}{\partial \lambda} &= -\frac{-\lambda + (1 - \lambda)}{\lambda^2} - \frac{1}{\lambda} = \frac{1 - \lambda}{\lambda^2} > 0. \end{aligned}$$

Using the implicit function theorem:

$$\frac{\partial t}{\partial \lambda} = -\frac{\frac{\partial F}{\partial \lambda}}{\frac{\partial F}{\partial t}} < 0.$$

and plugging these terms back into  $p$  yields:

$$\frac{\partial p}{\partial \lambda} = \frac{1}{\lambda} \frac{\partial t}{\partial \lambda} - \frac{1}{\lambda^2} (t - a) < 0.$$

When  $\lambda > 1$ ,  $p = t$ .  $t$  is not affected by  $\lambda$ . So  $\frac{\partial p}{\partial \lambda} = \frac{\partial t}{\partial \lambda} = 0$ .

#### A.4.7 Proof for Lemma 5 (Logit demand)

The logit demand is log-concave, since:

$$\begin{aligned} q'^2 - qq'' &= \beta^2 q^2 (1 - q)^2 - \beta^2 q^2 (1 - q)(1 - 2q) \\ &= \beta^2 q^2 (1 - q)(1 - q - 1 + 2q) = \beta^2 q^3 (1 - q) > 0. \end{aligned}$$

so that Assumptions 1 and 3 hold. In addition, Assumption 2 holds, since:

$$\frac{qq''}{q'^2} = \frac{\beta^2 q^2 (1-q)(1-2q)}{\beta^2 q^2 (1-q)^2} = \frac{1-2q}{1-q} = 1 - \frac{q}{1-q}$$

is decreasing in  $q$ , and thus increasing in  $p$ .

#### A.4.8 Proof for Lemma 6 (Exponential demand)

The exponential function is log-concave since:

$$q'^2 - qq'' = \beta^2 e^{-2\beta p} - \beta^2 e^{-\beta p} \cdot e^{-\beta p} = 0,$$

so that Assumptions 1 and 3 hold. In addition, Assumption 2 also holds, since:

$$(\log q)' + (\log q'')' - 2(\log(-q'))' = \beta + \beta - 2\beta = 0.$$

#### A.4.9 Proof for Lemma 7 ( $\rho$ -linear Demand)

The  $\rho$ -linear function is log-concave since:

$$q'^2 - qq'' = b^2 \frac{1}{\rho} (a - bp)^{2/\rho-2} > 0,$$

so that Assumptions 1 and 3 hold. In addition, Assumption 2 also holds, since:

$$(\log q)' + (\log q'')' - 2(\log(-q'))' = \frac{-b}{a - bp} \left( \frac{1}{\rho} + \frac{1}{\rho} - 2 - 2\left(\frac{1}{\rho} - 1\right) \right) = 0.$$

## Online Appendix B Experimental evidence on shopping behavior

Our experiment provided consumers with information on the availability of public pharmacies as an affordable alternative for purchasing drugs. This appendix studies whether consumers learned about the availability and attributes of public pharmacies, and whether knowing about them changed their shopping behavior in the short term. We estimate the equation:

$$y_i = \beta T_i + X_i' \gamma + \eta_{c(i)} + \varepsilon_i \quad (6)$$

where  $y_i$  is the outcome of interest;  $T_i$  indicates whether a consumer was treated;  $X_i$  is a vector of controls that includes the dependent variable at baseline along with consumer age, education, gender, and indicators for whether the consumer is covered by public insurance and whether a household member suffers a chronic condition;  $\eta_{c(i)}$  are county fixed effects. The coefficient  $\beta$  measures the average treatment effect of our informational intervention.

Information about public pharmacies rendered consumers more aware of their availability and attributes. Panel A in Table A.10 displays these results. Columns (1) and (2) show that information

increased awareness about the availability of the public pharmacy by 7 percentage points, from a baseline level of 77 percent. Moreover, columns (4) and (5) show that information shifted consumer perceptions about drug prices at public pharmacies, which is their most salient attribute. In particular, perceived public pharmacy prices decreased by 9 percent as a result of the intervention. We also find that perceived waiting time for receiving drugs at the public pharmacy increased, which is their main disadvantage relative to private pharmacies. In particular, perceived waiting time increased by 20 percent.<sup>42</sup> These results are consistent with consumers becoming aware of public pharmacies and their competitive advantages and disadvantages relative to private pharmacies as public pharmacies enter local markets.

Consumers also seem to have reacted to the intervention in terms of shopping behavior. Panel B in Table A.10 displays results from linear probability models for enrollment in the public pharmacy, the decision to purchase, and the plan to use the pharmacy in the future. Although estimates are imprecise, they are positive and economically meaningful. The point estimate in column (2) indicates a 2-percentage-points increase in enrollment with public pharmacies by treated households—almost a 30 percent increase relative to the mean of the control group. The results in column (5) imply a 2.3-percentage-points increase in purchases in public pharmacies by treated households—more than an 80 percent increase relative to a baseline share of 2.8 percent in the control group. Finally, column (8) shows that our intervention increased the extent to which households plan to use the public pharmacy by 5 percentage points, which is as much as 10 percent relative to the baseline level for the control group.

Households with members who suffer chronic conditions react more strongly to the treatment. Columns (3), (6), and (9) study heterogeneity along this margin. All effects are larger for households with chronic conditions, although the differences are not statistically significant. Moreover, the treatment effects on effective and planned purchases are marginally statistically significant for consumers with chronic conditions. Consumers with chronic conditions are more likely to periodically shop for drugs and thus the group for which short-term effects are more likely to be detectable. Moreover, in many cases, public pharmacies prioritize the provision of drugs to treat chronic conditions, and thus the information in our intervention may be less relevant for consumers without any household member with a chronic condition. Treatment effects on consumers without a household member with a chronic condition are indeed close to zero across outcomes.<sup>43</sup>

These results suggest that as public pharmacies enter local markets, consumers become aware of their entry, their relative advantages in terms of lower prices, and their relative disadvantages in terms of convenience. Moreover, our findings suggest that consumers value the availability of public pharmacies and some—particularly those affected by a chronic condition—substitute toward public pharmacies to take advantage of their lower drug prices.

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<sup>42</sup>We address concerns related to sample attrition by reporting bounds suggested by Lee (2009) in Table A.10-A. In all cases, point estimates for both the lower and upper bound have the same sign as our estimated treatment effects. However, in some cases, the point estimate of the bound is not statistically different from zero, which implies that under relatively negative attrition scenarios, our treatment effects are not distinguishable from zero.

<sup>43</sup>We report Lee bounds in Panel B in Table A.10 to address concerns about attrition. We find that point estimates for both the lower and upper bound for all outcomes have the same sign as our estimated treatment effects, although some of those bounds are not statistically different from zero.



## Online Appendix C The price effects of competition by public pharmacies

In this section, we develop a simple model of consumer choice and firm competition based on Chen and Riordan (2008). The goal is to illustrate the conditions under which the entry of an additional firm to a market induces an increase or a decrease in the prices set by an incumbent firm. The environment is simple but captures several features of our setting.

### C.1 Setup

**Environment.** There is a population of consumers of size one that faces the discrete choice problem of purchasing from the incumbent, purchasing from the entrant, or not purchasing at all, which is the outside option. We denote these options by  $j \in \{I, E, O\}$ , respectively. After normalizing the value of the outside option to 0, the value that consumer  $i$  gets from each option is

$$\begin{aligned} u_{iI} &= v_{iI} - p_r \\ u_{iE} &= v_{iE} - p_u \\ u_{iO} &= 0 \end{aligned}$$

where  $v_{ij}$  is the willingness to pay and  $p_j$  is the price of each option. Willingness to pay  $v_i$  is drawn from a differentiable joint distribution  $H(v)$ , and may feature average differences across firms, may be heterogeneous across consumers within each firm and may be correlated across firms. Consumers choose the option that gives the highest utility, so that the probability that consumer  $i$  chooses option  $j$  is

$$\sigma_{ij} = P(u_{ij} \geq u_{ik} \quad \forall k)$$

which induces demand functions

$$s_j = \int \sigma_{ij} h(v) dv$$

which naturally depend on the set of firms in the market.

On the supply side, the incumbent firm  $I$  chooses  $p^I$  to maximize profits  $s_I(p_I - c_I)$ , which leads to an optimal monopoly price  $p_I^m$  before entry and an optimal duopoly price  $p_I^d$  after entry. The entrant firm is meant to capture public pharmacies in our setting. As such, we assume it sets prices at marginal cost to satisfy a break-even condition, which is  $p_E^d = c_E$ .<sup>44</sup>

### C.2 When does entry increase prices?

The net price effects of entry depend on the relative importance of two competing forces: (i) the extent of substitution away from the monopolist, which imposes downward pressure on the incumbent price, and (ii) the extent to which the demand faced by the monopolist becomes steeper after entry, which imposes upward pressure on the incumbent price. To establish this intuition formally, we define  $F(v_I)$  as the marginal distribution of willingness to pay for the incumbent and  $G(v_E|v_I)$  as the distribution of willingness to pay for the entrant, conditional on that for the

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<sup>44</sup>All results hold for the case in which the entrant sets a profit-maximizing price.

incumbent. Both of these distributions are defined under the joint distribution  $H(v)$ . With this notation, we can restate Theorem 1 in Chen and Riordan (2008), which establishes that—under a few fairly general assumptions—the incumbent price will increase upon entry if and only if

$$\int_{p_I^m}^{\infty} [G(v|v) - G(p_I^m|v)]f(v)dv \leq (p_I^m - c_I) \int_{p_I^m}^{\infty} [g(p_I^m|v) - g(v|v)]f(v)dv$$

and will otherwise decrease.

This condition compares the magnitude of the two effects of entry. The left-hand side of the equation is the *market share effect* of entry. This term measures the difference between the market share the incumbent gets from charging the monopoly price as a monopoly and as a duopoly; that is, before and after entry. The more market share the entrant takes away from the incumbent, the stronger the incentives the incumbent has to decrease price in response to entry. The right-hand side of the equation is the *price sensitivity effect* of entry. The magnitude of this effect depends on the difference between the slope of the residual demand curve the incumbent faces before and after entry. The steeper the demand curve after entry relative to before entry, the lower the extent of substitution away from the incumbent from marginal consumers upon entry, and therefore the stronger the incentive of the incumbent to increase price upon entry.

The relative strength of these effects will largely depend on the distribution of consumer preferences. For example, the likelihood of a price increase is higher with a negative correlation in willingness to pay. In this case, substitution toward the entrant is lower than under a distribution of preferences with a positive correlation. Moreover, those who substitute away from the incumbent are consumers with a relatively low willingness to pay for the incumbent among those who purchase from the incumbent before entry, which leads to a steeper residual demand curve after entry.

### C.3 Simulation

In this section, we show the results of simulating the model. The goal is to show numerically how different parameter combinations yield different predictions regarding the sign of the price effect of entry.

**Specification.** A key input in the simulation is the joint distribution of willingness to pay for the firms in the market,  $H_v$ , which we assume follows a joint normal distribution:

$$\begin{pmatrix} v_I \\ v_E \end{pmatrix} \sim N \left( \begin{pmatrix} \delta_I \\ \delta_E \end{pmatrix}, \begin{pmatrix} \sigma_I^2 & \rho\sigma_I\sigma_E \\ \rho\sigma_I\sigma_E & \sigma_E^2 \end{pmatrix} \right)$$

where the mean willingness to pay for each firm is denoted by  $\delta_I$  and  $\delta_U$ . Differences between  $\delta_I$  and  $\delta_U$  capture vertical differentiation between firms and relative to the outside option. The dispersion of willingness to pay is captured by the variances  $\sigma_I^2$  and  $\sigma_E^2$ , and the correlation between the willingness to pay for the incumbent and the entrant is captured by  $\rho$ . If the willingness to pay is positively correlated ( $\rho > 0$ ), then consumers share similar preferences for both goods relative to the outside option. If instead willingness to pay is negatively correlated ( $\rho < 0$ ), then consumers with a strong taste for one of the firms have a weak taste for the other firm. This parameter

determines the extent to which the slope of demand the incumbent faces changes upon entry, which is key in determining the price effects of entry.

**Simulation details.** We simulate equilibrium prices and market shares for the environments before and after entry, for a range of parameters of the distribution of preferences. In particular, we set  $\delta_I$  and  $\delta_E$  so that  $(\delta_I + \delta_E)/2 = 10$  and  $\delta_I/\delta_E = k_\delta$  for a grid of values for  $k_\delta$  from 1 to 10; we set  $\sigma_I = \sigma_E = \sigma$  and construct a grid of values for  $\sigma$  from 1 to 15; and we construct a grid of values for  $\rho$  between -1 and 1. We set marginal costs at  $c_I = 6$  and  $c_E$ . For each combination of  $(k_\delta, \sigma, \rho)$ , we solve for optimal prices and resulting market shares before and after entry.

## C.4 Results

**Results on price effects and the distribution of preferences.** Our simulations illustrate that consumer preferences over firms play a key role in determining the equilibrium effects of entry on prices. Figure A.5 displays results for simulations over a grid of values for heterogeneity in preferences  $\sigma$  and correlation in preferences across firms  $\rho$ , for relative mean preferences of  $\delta_I/\delta_E = 4$ .

These results show two main patterns. First, the price charged by the incumbent firm is more likely to increase when preferences for the incumbent are more negatively correlated with those for the entrant. A more negative correlation implies that marginal consumers who substitute toward the entrant are those with a low willingness to pay for the incumbent, which makes the residual demand curve of the incumbent steeper and therefore imposes incentives to increase prices. This is consistent with a stronger price-sensitivity effect. Second, the results show that the price charged by the incumbent is more likely to increase when there is more dispersion in preferences, which is partly driven by the fact that when such dispersion is low, the demand curve is flatter and there is limited scope for price increases.

In the context of our setting and empirical results, this simulation suggests that the correlation between preferences for private and public pharmacies is likely negative. This suggests that pharmacy attributes—beyond drug prices—play an important role in pharmacy choice. An attribute that could be important in generating this pattern is heterogeneity in consumer locations relative to pharmacies: Consumers who live closer to private pharmacies are likely to pay more for them than for public pharmacies, whereas the opposite may be true for consumers who live closer to public pharmacies.

**Results on price effects and the relative quality of the entrant.** In addition to studying the conditions under which incumbent prices increase upon entry, we use the model to illustrate the importance of vertical quality difference in determining the penetration of the entrant and the differences in prices between the incumbent and the entrant. Figure A.6 shows results from simulations of the model for a grid of values for the relative quality of the incumbent  $\delta_I/\delta_E$ , while keeping average quality across firms fixed. We fix the remainder of the distribution of preferences to values such that the price of the incumbent increases; namely,  $\rho = -0.99$  and  $\sigma = 2.55$ .

We study the implications of vertical differentiation for market shares and prices. Panel A in Figure A.6 shows that while the entrant is able to steal market share from the incumbent, the extent of business stealing decreases substantially as the quality of the entrant relative to the incumbent decreases. Panel B in Figure A.6 shows that the incumbent price is higher when the quality of the entrant relative to the incumbent is lower. Furthermore, these results also show that the price effects of entry on the incumbent price depend on the relative quality of the entrant. The higher the

relative quality of the entrant, the more likely the incumbent price will decrease upon entry.

These results are consistent with our descriptive evidence and main empirical findings. In Section 3.1, we documented that public pharmacies entered the market offering lower quality along several dimensions, which suggests that  $\delta_I/\delta_E$  is relatively large in our setting. These results indeed imply that entrants with low relative quality have low penetration, allow the incumbent to sustain higher prices, and make it more likely that the incumbent will increase prices.

FIGURE A.1: EXAMPLES OF PRIVATE AND PUBLIC PHARMACIES



(a) Outside a private pharmacy



(b) Inside a private pharmacy



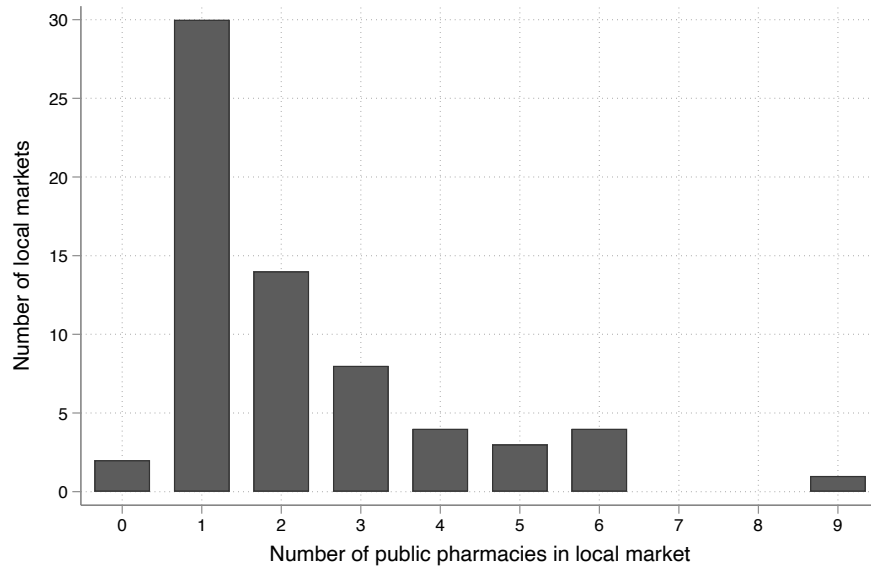
(c) Outside a public pharmacy



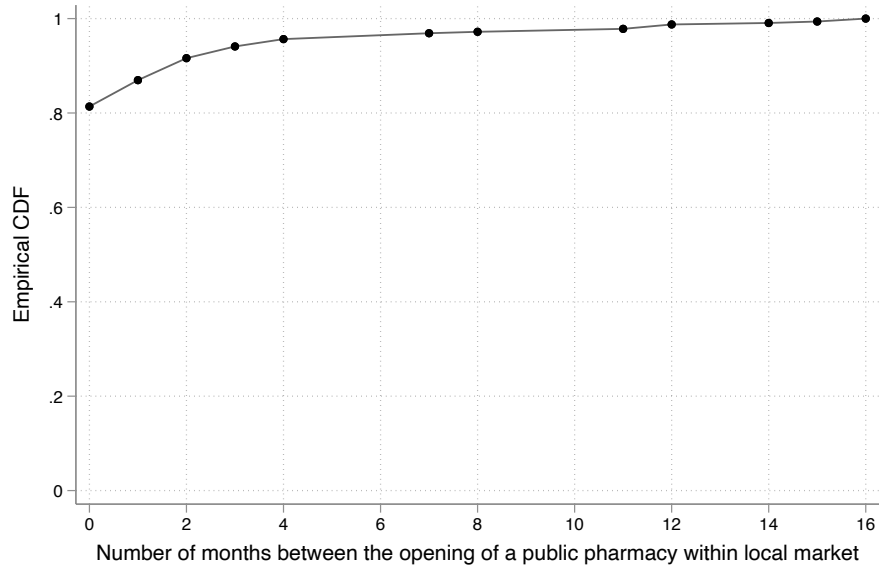
(d) Inside a public pharmacy

*Notes:* This figure displays photos of private and public pharmacies from the outside and inside. The private pharmacy in Panels (a) and (b) is a somewhat generic building and is one of the three leading chains in the market. The public pharmacy in Panels (c) and (d) is located in the city capital and is part of our experimental sample.

FIGURE A.2: NUMBER OF EVENTS PER MARKET AND TIME DISPERSION



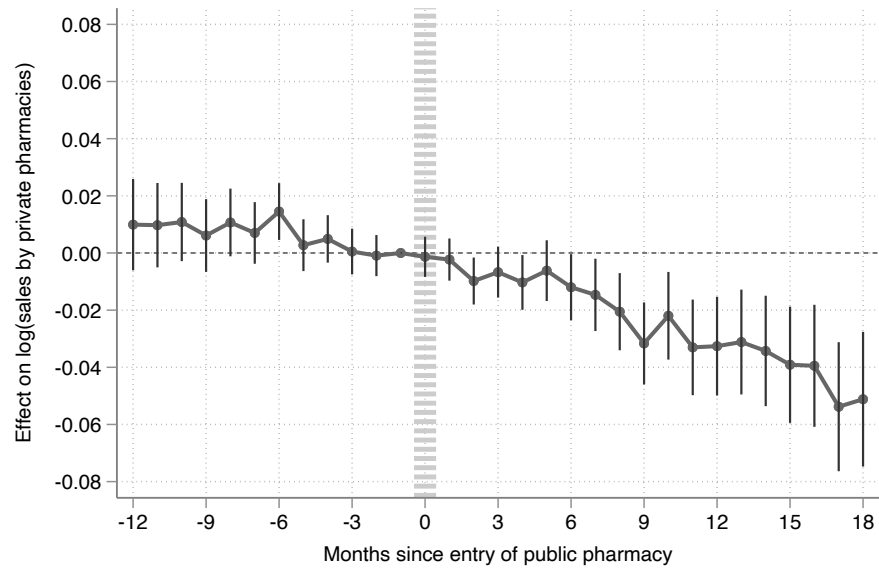
(a) Number of public pharmacy entries by market



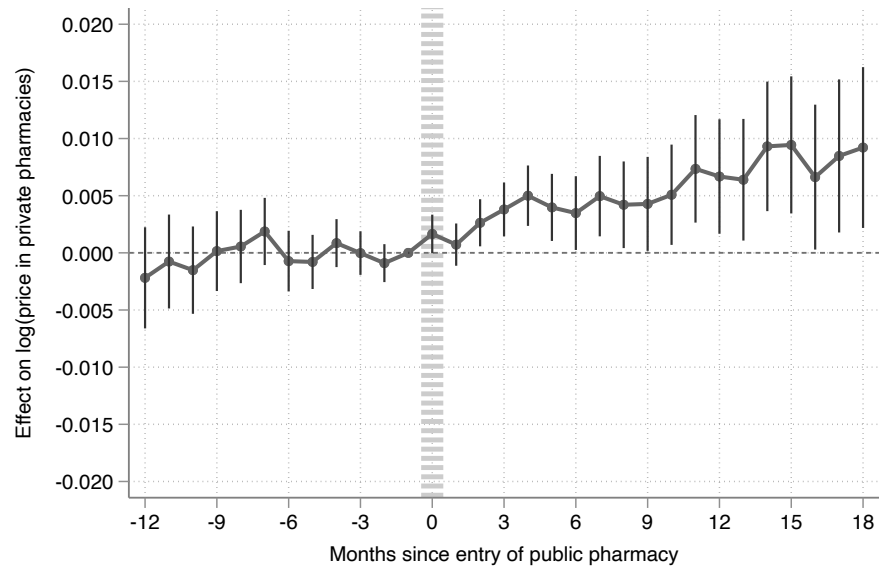
(b) Time between subsequent public pharmacy entries within a market

*Notes:* Panel (a) shows the distribution of the number of public pharmacies per local market. Panel (a) shows the cumulative distribution function of the dispersion of events within local markets. For example, more than 80 percent of events within the market occurred within the same month, which is by definition the case for markets with only one event.

FIGURE A.3: IMPACT OF PUBLIC PHARMACIES USING ONLY THE FIRST ENTRY IN A LOCAL MARKET



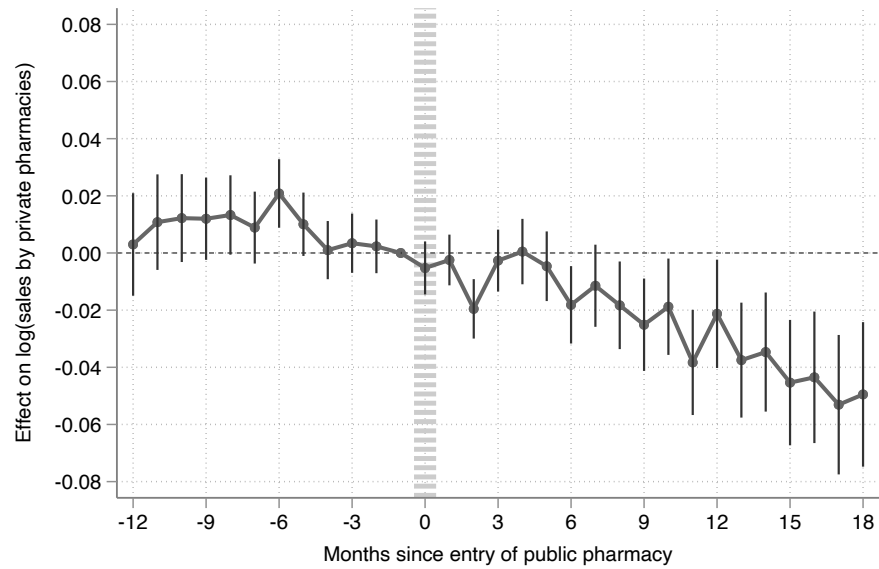
(a) Sales



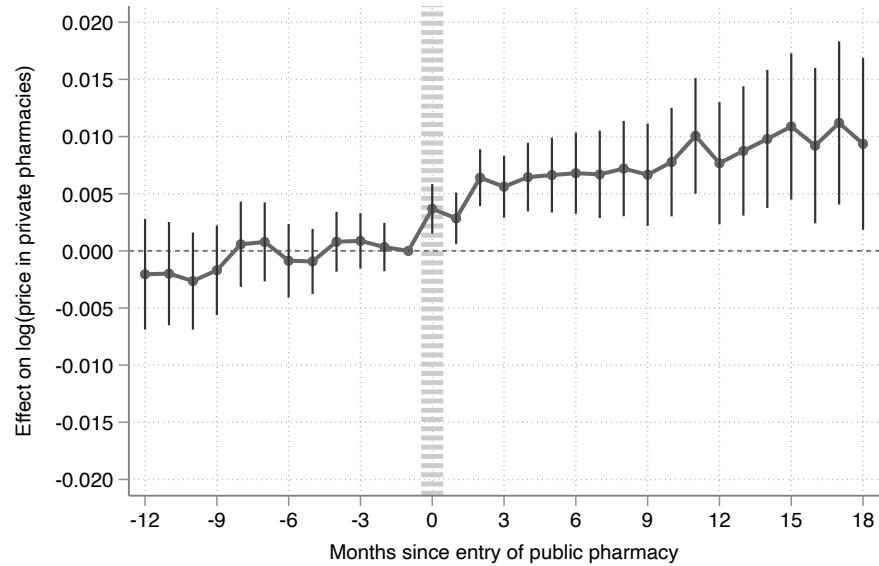
(b) Prices

*Notes:* These figures present event-study estimates of the impact of public pharmacies on private pharmacy sales in Panel (a), and on private pharmacy prices in Panel (b). The unit of observation is a molecule per market in a given month. The empirical strategy uses panel data for the period between 2014 and 2018 and exploits the staggered entry of public pharmacies from October 2015 onward in an event-study design. Treatment is defined as introduction of the *first* public pharmacy in the market. In Panel (a) the dependent variable is logged sales and in Panel (b) the dependent variable is logged prices. The *x*-axis indicates the month with respect to the opening of the public pharmacy, i.e., 18 means 18 months after the opening, and -12 means 12 months before the opening. Dots indicate estimated coefficients, and vertical lines indicate the corresponding 95 percent confidence intervals.

FIGURE A.4: IMPACT OF PUBLIC PHARMACIES AMONG MARKETS WITH EVENTS WITHIN LESS THAN 1 MONTH



(a) Sales

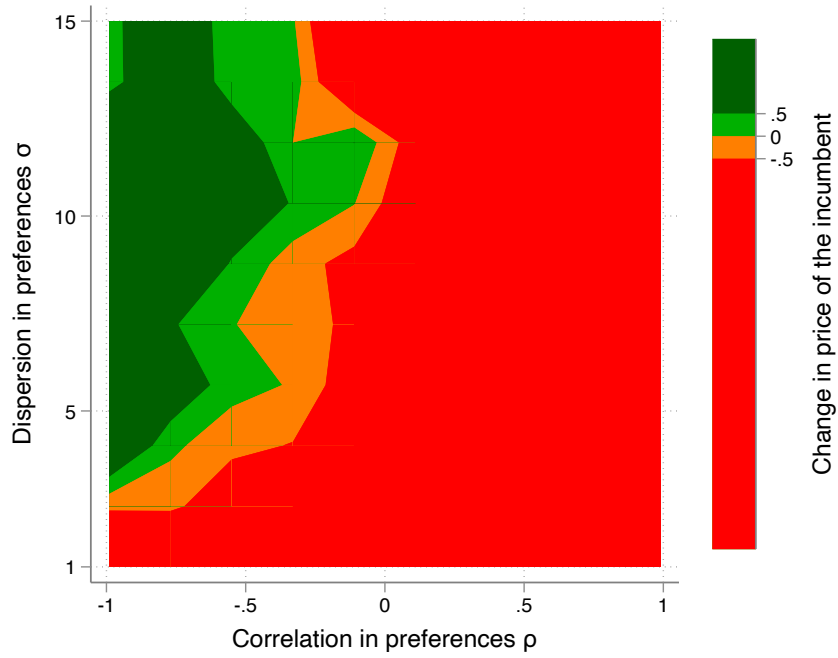


(b) Prices

*Notes:* These figures present event-study estimates of the impact of public pharmacies on private pharmacy sales in Panel (a), and on private pharmacy prices in Panel (b). The unit of observation is a molecule per market in a given month. The empirical strategy uses panel data for the period between 2014 and 2018 and exploits the staggered entry of public pharmacies from October 2015 onward in an event-study design. Treatment is defined as introduction of the *first* public pharmacy in the market. The sample only includes never-treated markets and markets with either one event or in which events are no more than 1 month apart. In Panel (a) the dependent variable is logged sales and in Panel (b) the dependent variable is logged prices. The *x*-axis indicates the month with respect to the opening of the public pharmacy, i.e., 18 means 18 months after the opening, and -12 means 12 months before the opening. Dots indicate estimated coefficients and vertical lines indicate the corresponding 95 percent confidence intervals.

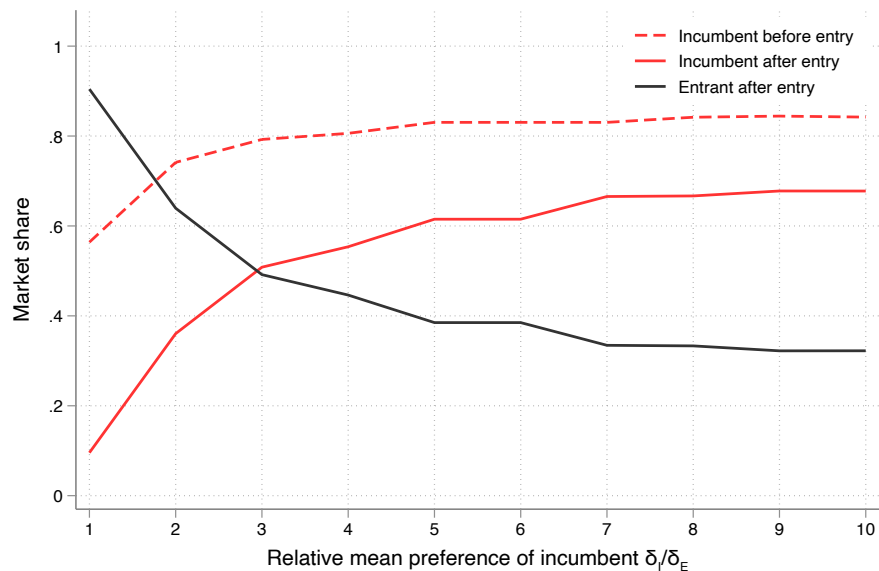


FIGURE A.5: SIMULATIONS FOR THE PRICE EFFECTS OF ENTRY

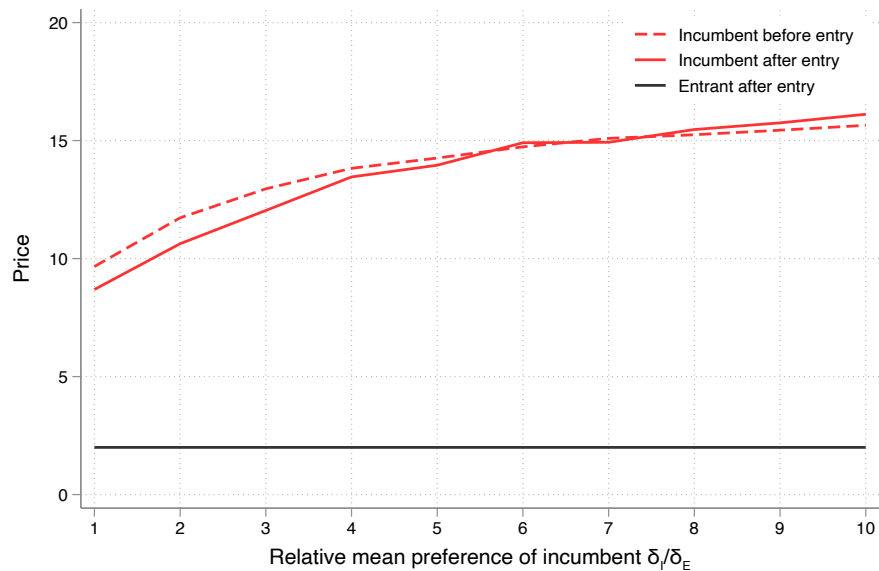


*Notes:* This figure plots simulated effects of entry on the price the incumbent charges, as discussed in Online Appendix C. The plot provides results for a grid of values of  $\sigma$  and  $\rho$ , under mean preferences for the incumbent and entrant  $\delta_I/\delta_E = 4$ , although the results are qualitatively similar for different values of the latter. The red region indicates that the incumbent price *decreases*, whereas the green region indicates that the incumbent price *increases* for each distributions of preferences, respectively.

FIGURE A.6: SIMULATIONS FOR THE ROLE OF RELATIVE QUALITY IN EQUILIBRIUM OUTCOMES



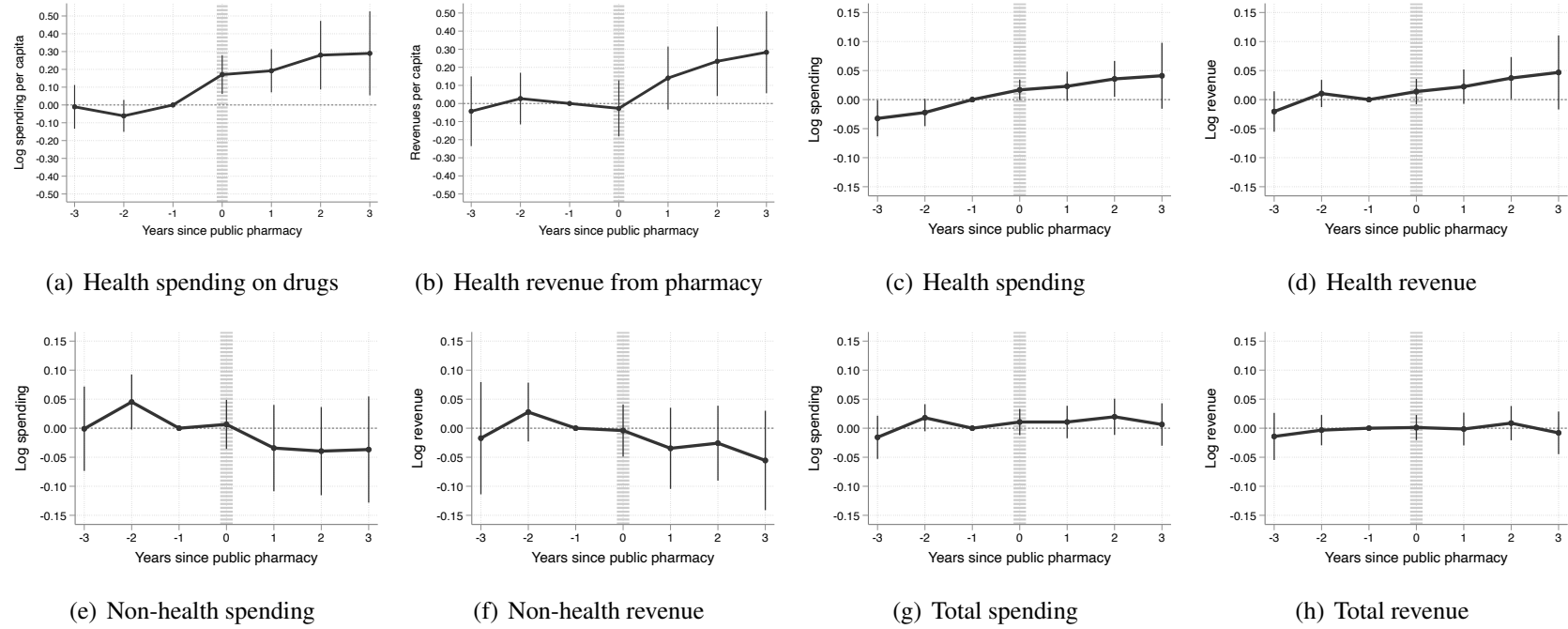
(a) Equilibrium market shares



(b) Equilibrium prices

*Notes:* Both panels display equilibrium outcomes for the incumbent and entrant, before and after entry for a range of values for relative quality of the incumbent  $\delta_I/\delta_E$ , while keeping the average quality of both firms fixed. Panel (a) displays equilibrium market shares, whereas Panel (b) displays equilibrium prices. Incumbent outcomes are plotted in red, while entrant outcomes are plotted in black. Outcomes before entry are plotted in dashed lines, while outcomes after entry are plotted in dashed lines.

FIGURE A.7: EVENT STUDY ESTIMATES FOR EFFECTS ON MUNICIPAL FINANCE

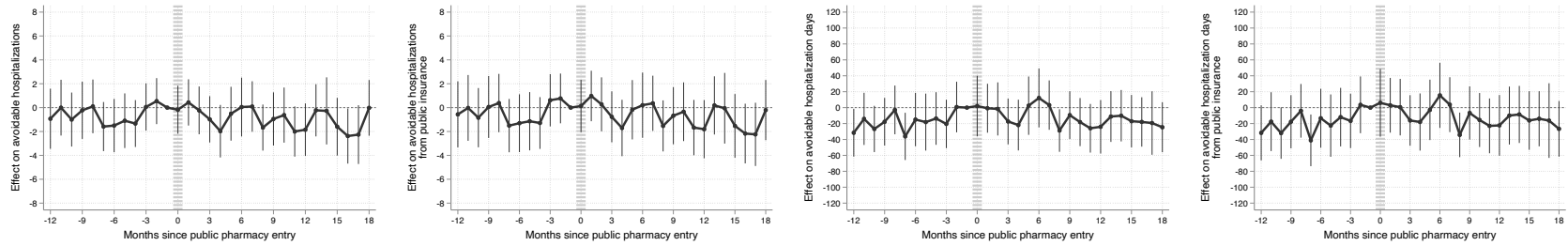


*Notes:* These figures present event study estimates for the impact of public pharmacies on municipal finance using panel data for 2013-2019. Municipal spending and revenue are measured in monetary units per capita. Each plot displays results from an event study version of equation (3) given by:

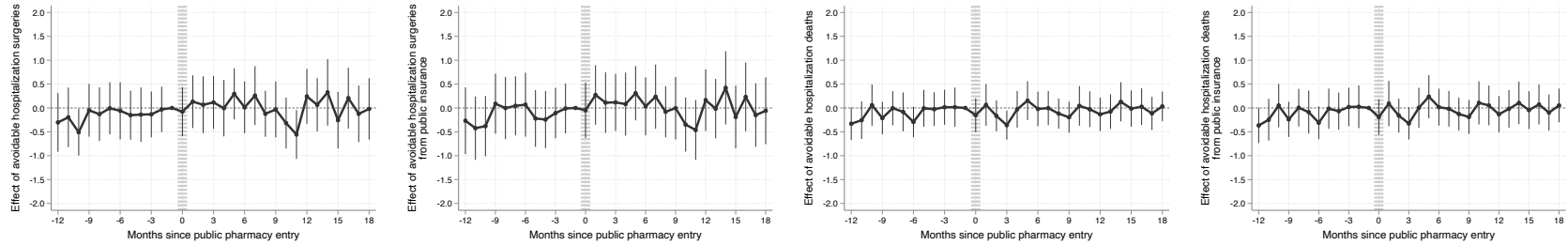
$$y_{ct} = \sum_{k=-3}^3 \delta_k D_{ct}^k + \theta_c + \lambda_t + \varepsilon_{ct},$$

where the outcomes are the measures of municipal finance (revenue, spending) in logarithms and treatment dummies are defined with respect to the first year with a public pharmacy. All regressions include county fixed effects  $\theta_c$  and year fixed effects  $\lambda_t$ . Dots indicate point estimates and vertical lines indicate the corresponding 95 percent confidence intervals.

FIGURE A.8: EVENT STUDY ESTIMATES FOR EFFECTS ON AVOIDABLE HOSPITALIZATIONS



(a) Number of hospitalizations, all insurance (b) Number of hospitalizations, public insurance (c) Days of hospitalizations, all insurance (d) Days of hospitalizations, public insurance



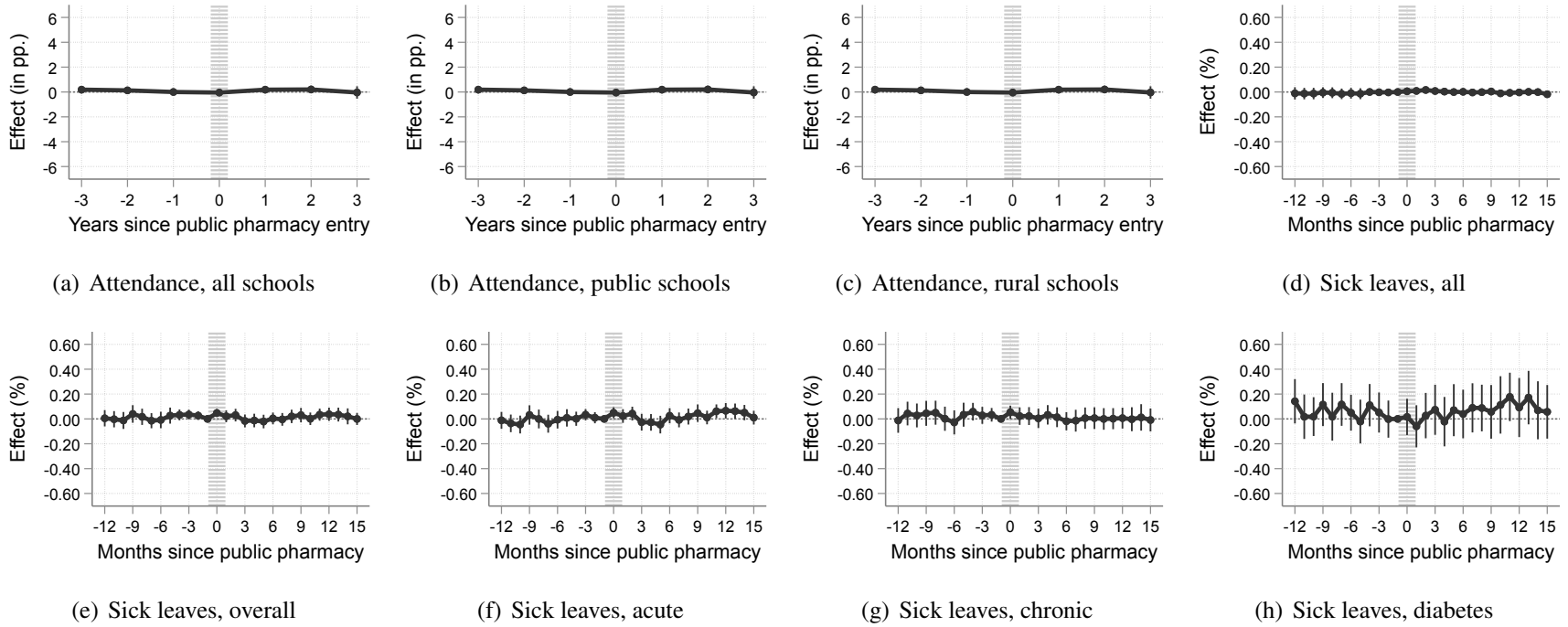
(e) Number of surgeries, all insurance (f) Number of surgeries, public insurance (g) Number of deaths, all insurance (h) Number of deaths, public insurance

Notes: Each plot displays results from an event study version of equation (3) given by:

$$y_{ct} = \sum_{k=-12}^{18} \delta_k D_{ct}^k + \theta_c + \lambda_t + \varepsilon_{ct},$$

where the outcomes are the same measures of avoidable hospitalization events as in Table 4 and treatment dummies  $D_{ct}^k$  are defined as a month  $t$  which is exactly  $k$  months after event time in county  $c$ . We normalize  $\delta_{k=-1} = 0$ , so we interpret all coefficients  $\delta_k$  as the effect of a public pharmacy's opening on the dependent variable exactly  $k$  months after its entry. Dots indicate point estimates and vertical lines indicate the corresponding 95 percent confidence intervals.

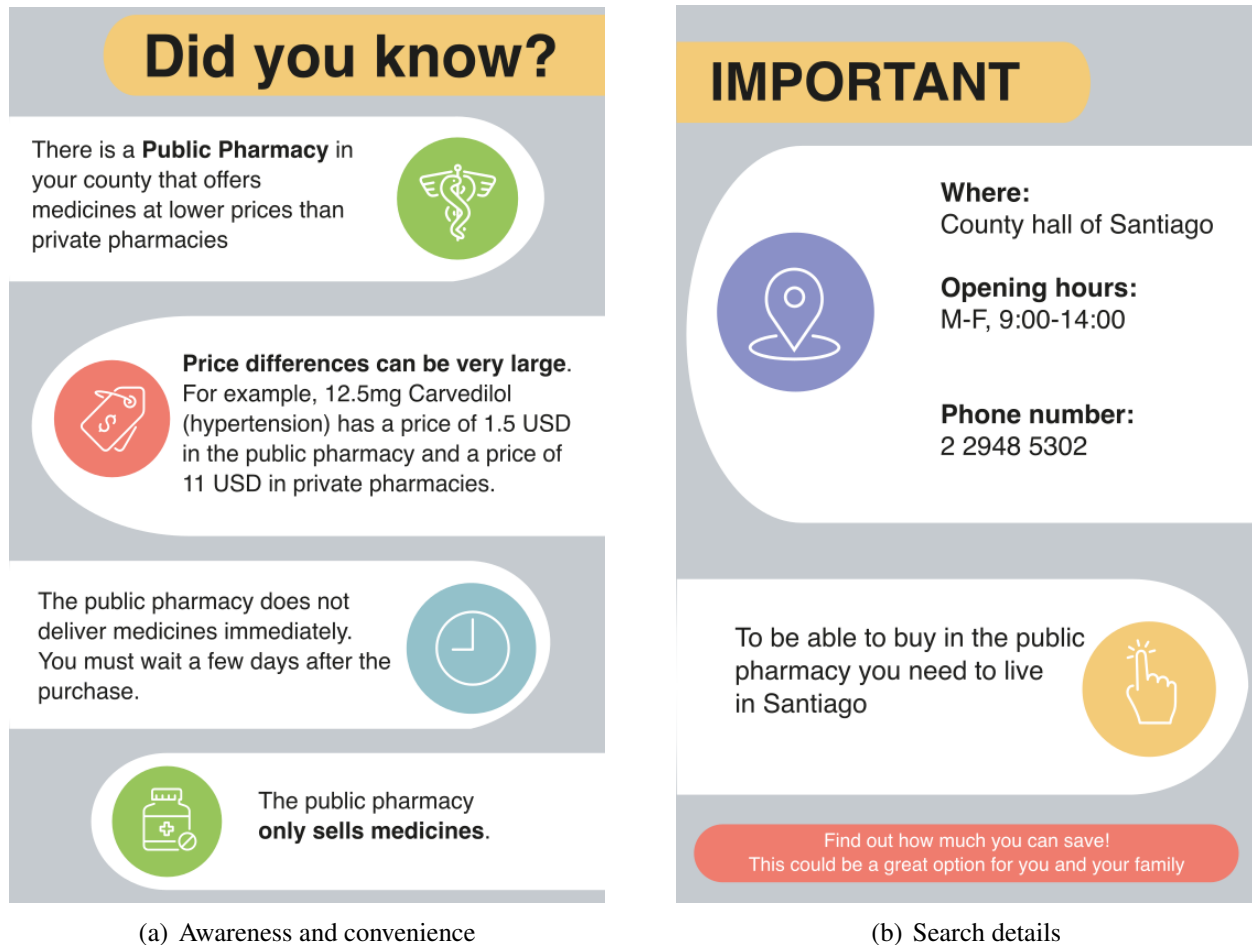
FIGURE A.9: OTHER HEALTH OUTCOMES, EVENT STUDY EVIDENCE



$$y_{ct} = \sum_{k=\kappa_1}^{\kappa_2} \delta_k D_{ct}^k + \theta_c + \lambda_t + \varepsilon_{ct}$$

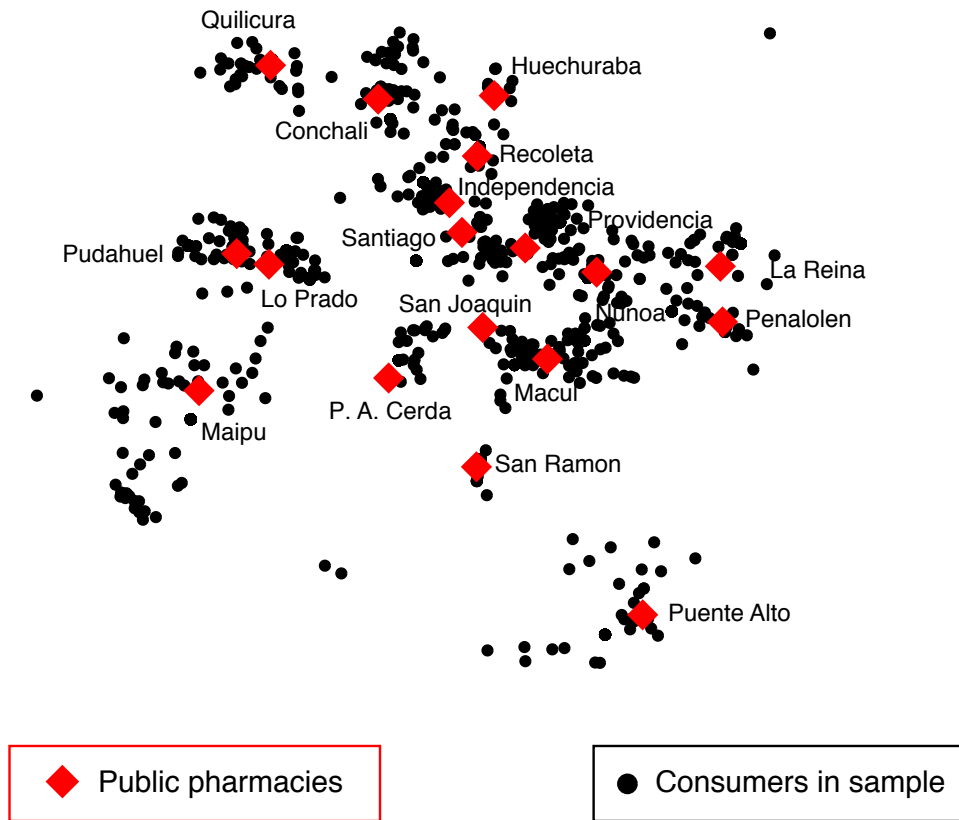
where the outcomes are school attendance in percentages ( $\in [0, 100]$ ) and the number of sick leave per capita. Treatment indicators are defined with respect to the first year (panels a-c) or month (panels d-h) with a public pharmacy. All estimates include county fixed effects  $\theta_c$  and year (or month-year) fixed effects  $\lambda_t$ . Each dot represents a coefficient and vertical lines indicate 95 percent confidence intervals.

FIGURE A.10: INFORMATIONAL TREATMENT



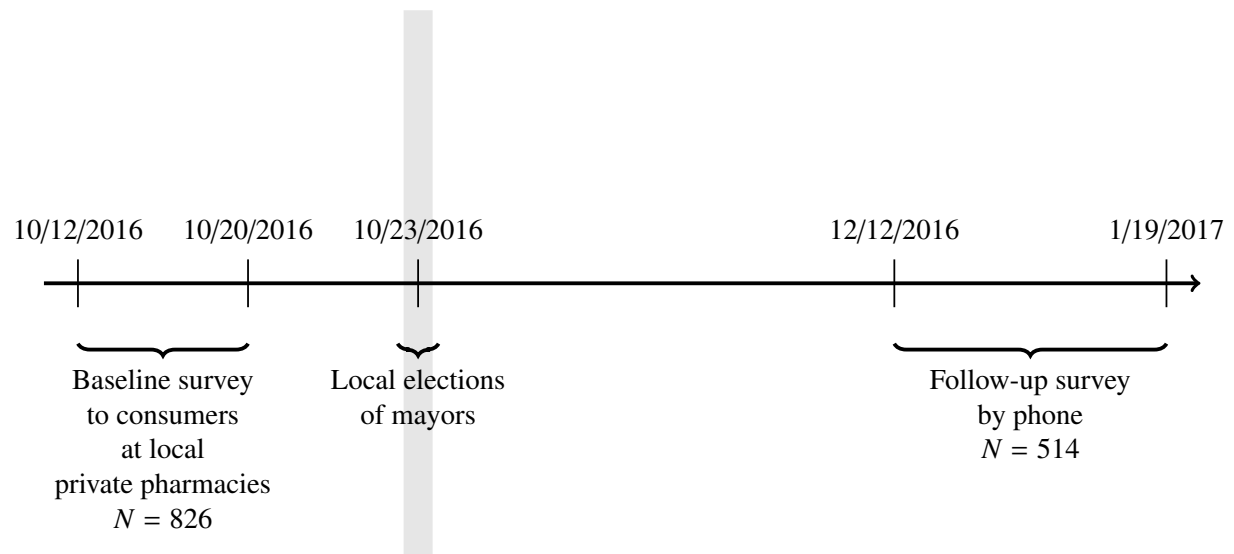
*Notes:* This figure displays the informational interventions delivered as part of the field experiment. Panel (a) displays the first part of the treatment, which aimed to increase awareness of the public pharmacy. It introduces the public pharmacy and states that it offers lower prices than private pharmacies and that it may take longer to deliver products. Panel (b) displays the second part, which aim to reduce search costs for participants by including detailed location and contact information for the public pharmacy, hours of operation, and eligibility requirements, tailored to the county of each participant.

FIGURE A.11: LOCATION OF PHARMACIES AND CONSUMERS IN EXPERIMENTAL SAMPLE



*Notes:* This figure displays the location of public pharmacies and consumers surveyed in the context of the field experiment. We surveyed 826 people at baseline outside randomly selected private pharmacies located in 18 counties within the city capital. All of these counties had a public pharmacy at the time of the baseline survey.

FIGURE A.12: TIMELINE OF EXPERIMENT EVENTS



*Notes:* This timeline displays the main events in our field experiment. Baseline surveys were implemented outside randomly chosen private pharmacies in counties with a public pharmacy. Local elections are a single-day election held every 4 years in which citizens in all 344 counties vote for a mayor using simple majority rule. Follow-up surveys were implemented during a 1-month period to minimize attrition.



TABLE A.1: WITHIN-COUNTY ANALYSIS OF PUBLIC PHARMACY ENTRY

	(1)	(2)	(3)	(4)
	1(Public pharmacy)			
Private pharmacies in 2014	0.022*** (0.004)	0.032*** (0.006)	0.030*** (0.005)	0.026*** (0.004)
Schools in 2014	0.016*** (0.002)	0.011*** (0.002)	0.011*** (0.002)	0.007*** (0.001)
Cell size is (in meters):	1,000	800	600	400
Cells	10,167	14,046	21,885	43,695
Mean of dependent variable	0.014	0.010	0.006	0.003
Mean of private pharmacies	0.198	0.142	0.088	0.049
County fixed effects	Yes	Yes	Yes	Yes

*Notes:* The unit of observation is a geographic cell within a county. We use all 146 counties with a public pharmacy operating by December 2018. Private pharmacies are measured in the year 2014, before the opening of public pharmacies. The estimating sample restricts attention to “populated cells,” i.e., cells within the convex hull of schools in 2014. Different columns display results for different definitions of cell size, from 1,000×1,000 meters in column (1) to 400×400 meters in column (4). Standard errors clustered by county.

TABLE A.2: EFFECT ON DRUG SALES AND PRICES IN THE PRIVATE MARKET

	(1) log(sales)	(2) log(price)
Panel A: Main estimates		
All sample ( $\beta^{\text{jump}}$ )	0.0071* (0.0043) [0.0074]	0.0033*** (0.0011) [0.0022]
All sample ( $\beta^{\text{phase in}}$ )	-0.0029*** (0.0005) [0.0008]	0.0005*** (0.0001) [0.0006]
R-squared	0.54	0.85
Panel B: Heterogeneity by chronic condition		
Molecules for chronic conditions ( $\beta^{\text{jump}}_{\text{chronic}}$ )	-0.0036 (0.0050) [0.0082]	0.0029** (0.0013) [0.0023]
Molecules for non-chronic conditions ( $\beta^{\text{jump}}_{\text{non-chronic}}$ )	0.0219*** (0.0079) [0.0103]	0.0037* (0.0020) [0.0027]
Molecules for chronic conditions ( $\beta^{\text{phase in}}_{\text{chronic}}$ )	-0.0028*** (0.0005) [0.0008]	0.0005*** (0.0001) [0.0006]
Molecules for non-chronic conditions ( $\beta^{\text{phase in}}_{\text{non-chronic}}$ )	-0.0030*** (0.0007) [0.0010]	0.0005*** (0.0002) [0.0007]
Panel C: Heterogeneity by relative public/private product variety		
High public-private variety ratio ( $\beta^{\text{jump}}_{\text{high variety}}$ )	-0.0110** (0.0048) [0.0096]	0.0037*** (0.0012) [0.0025]
Low public-private variety ratio ( $\beta^{\text{jump}}_{\text{low variety}}$ )	0.0039 (0.0057) [0.0080]	0.0031** (0.0015) [0.0023]
High public-private variety ratio ( $\beta^{\text{phase in}}_{\text{high variety}}$ )	-0.0035*** (0.0005) [0.0010]	0.0007*** (0.0001) [0.0006]
Low public-private variety ratio ( $\beta^{\text{phase in}}_{\text{low variety}}$ )	-0.0024*** (0.0006) [0.0009]	0.0003** (0.0002) [0.0006]
Panel D: Heterogeneity by distance to private pharmacy		
Private pharmacies are close to public pharmacy ( $\beta^{\text{jump}}_{\text{close}}$ )	0.0074 (0.0053) [0.0083]	0.0041*** (0.0013) [0.0024]
Private pharmacies are far from public pharmacy ( $\beta^{\text{jump}}_{\text{far}}$ )	0.0068 (0.0052) [0.0103]	0.0024 (0.0014) [0.0027]
Private pharmacies are close to public pharmacy ( $\beta^{\text{phase in}}_{\text{close}}$ )	-0.0033*** (0.0006) [0.0010]	0.0002* (0.0002) [0.0007]
Private pharmacies are far from public pharmacy ( $\beta^{\text{phase in}}_{\text{far}}$ )	-0.0025*** (0.0006) [0.0009]	0.0008*** (0.0002) [0.0006]
Observations	681,120	649,885
Molecule-by-month FE	Yes	Yes
Molecule-by-market FE	Yes	Yes

*Notes:* This table presents our parametric estimates of the effects of public pharmacy entry on private market outcomes. We estimate the parameters  $\beta^{\text{jump}}$  and  $\beta^{\text{phase in}}$  using an exposure difference-in-differences design that leverages the staggered introduction of public pharmacies in the panel data of molecules observed by market and month in the period 2014-2018. The parameter  $\pi^{\text{jump}}$  measures the immediate impact of public pharmacies and  $\pi^{\text{phase in}}$  the additional impact by each year of operation. In Panel B, exposure to public pharmacies is interacted with an indicator for whether a molecule is targeted toward a chronic condition or not. In Panel C, exposure is interacted with an indicator for whether there is a high ratio of variety of products within a molecule in public pharmacies relative to private pharmacies. In Panel D, exposure is interacted with an indicator for whether in the local market private pharmacies were located relatively close to the public pharmacy. Standard errors clustered at the molecule-by-market level displayed in parentheses. \*\*\* p<0.01, \*\* p<0.05, \* p<0.1. We also provide standard errors clustered at the local market level and are displayed in square brackets.

TABLE A.3: MUNICIPAL FINANCE, FULL REGRESSION COEFFICIENTS

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Sub-categories of health related to public pharmacies		All health services		Non-health services		All services	
	Spending	Revenue	Spending	Revenue	Spending	Revenue	Spending	Revenue
$\pi^{\text{jump}}$ : Public pharmacy	0.129 (0.084)	-0.139 (0.104)	0.023* (0.013)	0.004 (0.013)	0.011 (0.036)	-0.003 (0.038)	0.004 (0.017)	0.001 (0.016)
$\pi^{\text{phase in}}$ : Public pharmacy $\times$ trend	0.055 (0.046)	0.130*** (0.043)	0.010 (0.006)	0.012* (0.006)	-0.023 (0.017)	-0.011 (0.017)	0.005 (0.007)	0.004 (0.007)
Avg. dep. var. in 2014	9.144	6.518	182.60	181.43	513.08	548.73	695.68	730.15
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Counties	320	320	321	321	322	322	322	322
Observations	2,200	2,205	2,240	2,240	2,228	2,227	2,243	2,243

*Notes:* This table presents our estimates for the impact of public pharmacies on municipal finances. We observe a panel of counties every year in the period 2013–2019 and exploit the staggered entry of pharmacies in a parametric event study analysis. The dependent variable is the logarithm of total spending (in U.S. dollars) per capita (2013 population) in odd columns and the logarithm of total revenue per capita in even columns. The parameter  $\pi^{\text{jump}}$  measures the immediate impact of public pharmacies and  $\pi^{\text{phase in}}$  the additional impact by each year of operation. We focus on 18-month effects to compare the cost of public pharmacies with their impact on sales and prices in private pharmacies (Panel (a) of Table 2). Standard errors clustered at the county level are displayed in parentheses. Significance level: \*\*\* p<0.01, \*\* p<0.05, \* p<0.1.

TABLE A.4: EFFECTS OF PUBLIC PHARMACIES ON MUNICIPAL FINANCE FOR ALTERNATIVE TRANSFORMATIONS OF THE DEPENDENT VARIABLE

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Sub-categories of health related to public pharmacies		All health services		Non-health services		All services	
	Spending	Revenue	Spending	Revenue	Spending	Revenue	Spending	Revenue
<b>Panel A: asinh(y)</b>								
Public pharmacy 18-month effect	0.289*** (0.055)	0.219*** (0.065)	0.047*** (0.015)	0.033** (0.016)	0.002 (0.080)	-0.015 (0.076)	0.015 (0.015)	0.010 (0.015)
<b>Panel B: log(y+1)</b>								
Public pharmacy 18-month effect	0.232*** (0.044)	0.163*** (0.053)	0.044*** (0.014)	0.030** (0.015)	-0.047 (0.034)	-0.035 (0.032)	0.015 (0.015)	0.010 (0.015)
<b>Panel C: log(y+0.001)</b>								
Public pharmacy 18-month effect	0.297** (0.132)	0.305** (0.134)	0.048*** (0.016)	0.034** (0.016)	-0.048 (0.035)	-0.030 (0.034)	0.015 (0.015)	0.010 (0.014)
<b>Panel D: log(y+10)</b>								
Public pharmacy 18-month effect	0.078*** (0.017)	0.039* (0.021)	0.027** (0.013)	0.015 (0.013)	-0.044 (0.030)	-0.032 (0.031)	0.010 (0.014)	0.006 (0.014)
Avg. dep. var. in 2014	8.945	6.376	170.36	167.09	525.32	563.07	695.68	730.15
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Counties	322	322	321	321	322	322	321	322
Observations	2,248	2,248	2,243	2,243	2,228	2,228	2,243	2,243

*Notes:* This table presents our estimates for the impact of public pharmacies on municipal finances. We observe a panel of counties every year in the period 2013–2019 and exploit the staggered entry of pharmacies in a parametric event study analysis. The dependent variable is some transformation (see panel header) of total spending per capita (2013 population) in odd columns and the same transformation of total revenue per capita in even columns. The 18-month effect is the linear combination of regression coefficients  $\pi^{\text{jump}} + (1.5 + 1) \times \pi^{\text{phase in}}$ . We focus on 18-month effects to compare the cost of public pharmacies with their impact on sales and prices in private pharmacies (Panel (a) of Table 2). Standard errors clustered at the county level are displayed in parentheses. Significance level: \*\*\* p<0.01, \*\* p<0.05, \* p<0.1.

TABLE A.5: EFFECT ON AVOIDABLE HOSPITALIZATIONS ASSOCIATED WITH CHRONIC DISEASES, FULL REGRESSION COEFFICIENTS

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Avoidable hospitalizations per 100,000 inhabitants							
	Number of hospitalizations		Days of hospitalizations		Number of surgeries		Number of deaths	
$\pi^{\text{jump}}$ : Public pharmacy	0.332 (0.489)	0.586 (0.532)	13.897* (7.091)	17.457** (7.838)	0.187 (0.149)	0.235 (0.166)	-0.035 (0.069)	-0.015 (0.077)
$\pi^{\text{phase in}}$ : Public pharmacy $\times$ trend	-0.064 (0.049)	-0.085 (0.052)	-1.039* (0.578)	-1.185* (0.646)	-0.004 (0.013)	-0.008 (0.014)	0.007 (0.006)	0.007 (0.006)
Health insurance	All	Public	All	Public	All	Public	All	Public
Mean of dep. var. in 2014	17.93	19.18	158.1	172.5	1.724	1.907	0.736	0.828
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Month fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
$R^2$	0.450	0.664	0.264	0.602	0.139	0.573	0.062	0.598
Counties	344	344	344	344	344	344	344	344
Observations (county-month-years)	28,320	28,320	28,320	28,320	28,320	28,320	28,320	28,320

*Notes:* This table presents our estimates for the impact of public pharmacies on avoidable health outcomes. The outcomes of interest are the number of hospitalizations (columns 1-2), days of hospitalizations (3-4), number of surgeries (columns 5-6), and number of deaths (columns 7-8). For each outcome, the first column uses the count of the outcome per 100,000 inhabitants in a county regardless of individual health insurance, and the second column restricts that count to individuals with publicly provided insurance (FONASA). We observe a panel of counties every month in the period 2013–2019 and exploit the staggered entry of pharmacies in a parametric event study analysis. The parameter  $\pi^{\text{jump}}$  measures the immediate impact of public pharmacies and  $\pi^{\text{phase in}}$  the additional impact by each year of operation. We report the mean of the dependent variable for 2014 among counties that ever introduce a public pharmacy, the year before most public pharmacies entered the market. Standard errors clustered at the county level are displayed in parentheses. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

TABLE A.6: PUBLIC PHARMACIES AND OTHER HEALTH OUTCOMES

	School attendance (county-year panel)			Sick leave (county-month panel)				
	All schools	Public schools	Rural schools	All diseases	Overall	Acute	Chronic	Diabetes
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Public pharmacy 18-month effect	0.137 (0.131)	0.335 (0.209)	-0.081 (0.151)	0.002 (0.023)	-0.001 (0.032)	0.022 (0.032)	-0.011 (0.053)	-0.017 (0.126)
Observations	2,070	2,031	1,802	20,002	18,194	17,441	15,096	8,509
Counties	345	345	301	346	341	336	336	311
County fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Avg. dependent variable	90.9	89.8	93.5	7.19	3.69	3.30	2.77	1.56

*Notes:* This table present difference-in-differences estimates for the impact of public pharmacies on school attendance using annual panel data for 2014-2019 and on sick leave using monthly panel data for 2015-2019. The former is administrative data from Ministry of Education and the latter is administrative data from the funding branch of the Ministry of Health (FONASA). Each column displays results from an event study regression given by:

$$y_{ct} = \theta_c + \lambda_t + \pi^{\text{jump}} PP_{ct} + \pi^{\text{phase in}} PP_{ct}(t - t_e^* + 1) + \varepsilon_{ct},$$

where the outcomes are school attendance in percentages ( $\in [0, 100]$ ) and the number of sick leave per capita in county  $c$  and year  $t$ ,  $PP_{ct}$  indicates the share of the year with a public pharmacy in county  $c$ , and  $(t - t_e^*)$  measures the number of years since the opening of the public pharmacy. All regressions include county fixed effects  $\theta_c$  and year (or month-year) fixed effects  $\lambda_t$ . Columns (1)-(3) use school absenteeism as dependent variable (years 2014-2019). Columns (4) and (5) use the logarithm of the total number of sick leave per 100,000 inhabitants (years 2015-2019). Standard errors are clustered at the county level. Significance level: \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

TABLE A.7: BALANCE IN COVARIATES ACROSS ATTRITION STATUS

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Non-Attriters vs Attriters			Panel B: Non-Attriters		
Variable	Non-Attriters	Attriters	$p$ -value $H_0 : (1) = (2)$	Control	Treatment	$p$ -value $H_0 : (4) = (5)$
Monthly drug expenditure	75.44 (71.93)	78.48 (70.37)	0.57	78.05 (75.50)	73.56 (69.31)	0.54
Chronic condition in household	0.61 (0.49)	0.49 (0.50)	0.00	0.61 (0.49)	0.61 (0.49)	0.65
Age	46.70 (16.67)	44.60 (18.08)	0.09	46.62 (16.84)	46.77 (16.57)	0.62
Education higher than HS	0.53 (0.50)	0.52 (0.50)	0.89	0.54 (0.50)	0.52 (0.50)	0.72
Female	0.64 (0.48)	0.58 (0.49)	0.06	0.62 (0.49)	0.66 (0.47)	0.74
Public insurance	0.63 (0.48)	0.66 (0.47)	0.34	0.62 (0.49)	0.63 (0.48)	0.31
Day with internet (1-7)	5.26 (2.84)	5.43 (2.71)	0.40	5.12 (2.92)	5.35 (2.78)	0.37
Day with social media (1-7)	5.22 (2.89)	5.34 (2.82)	0.56	5.07 (2.96)	5.32 (2.83)	0.17
Employed	0.63 (0.48)	0.64 (0.48)	0.74	0.59 (0.49)	0.65 (0.48)	0.82
Supports incumbent	0.48 (0.50)	0.56 (0.50)	0.09	0.50 (0.50)	0.47 (0.50)	0.23
Voted in previous election	0.76 (0.43)	0.70 (0.46)	0.06	0.74 (0.44)	0.78 (0.41)	0.88
Knows public pharmacy	0.67 (0.47)	0.60 (0.49)	0.04	0.64 (0.48)	0.69 (0.46)	0.08
Perceived relative price of public pharmacy	0.46 (0.23)	0.47 (0.18)	0.54	0.46 (0.18)	0.46 (0.26)	0.55
Perceived days to delivery at private pharmacy	8.52 (12.00)	8.53 (12.73)	1.00	9.71 (14.74)	7.67 (9.49)	0.80
Observations	514	312		216	298	

*Notes:* Columns (1) and (2) display the mean and standard deviation of different covariates at baseline for sample non-attriters and attriters, respectively. Column (3) displays the p-value from a test of equality of means across both groups. Columns (4) and (5) display the mean and standard deviation of different covariates at baseline for treatment and control group within the group of non-attriters surveyed at follow-up. Column (6) displays the p-value from a test of equality of means across both groups within the group of non-attriters surveyed at follow-up.

TABLE A.8: WAS A TREATMENT DELIVERED?

	(1)	(2)	(3)	(4)
	Delivered	Explained	Content	Useful
Treatment	0.107*** (0.033)	0.238*** (0.043)	0.304*** (0.059)	0.624 (0.438)
Constant	0.769*** (0.025)	0.440*** (0.033)	0.379*** (0.049)	7.208*** (0.379)
Observations	514	514	297	191
R-squared	0.020	0.060	0.083	0.011

*Notes:* This table displays results from different regressions of measures of treatment delivery on indicators for each of the treatment groups. Column (1) uses an indicator for treatment delivery as an outcome; column (2) uses an indicator for a treatment's being explained; column (3) uses an indicator for whether the participant recalls that the treatment was related to public pharmacies, conditional on receiving it; and column (4) uses a response on a scale from 1 to 10 regarding the usefulness of information, conditional on recalling the content. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .



TABLE A.9: BALANCE IN COVARIATES BETWEEN TREATMENT AND CONTROL GROUP

Variable	(1)	(2)	(3)
	Control	Treatment	$p$ -value $H_0 : (1) = (2)$
Monthly drug expenditure	76.31 (73.54)	76.69 (69.97)	0.94
Chronic condition in household	0.57 (0.50)	0.56 (0.50)	0.84
Age	45.25 (16.81)	46.32 (17.50)	0.39
Education higher than HS	0.54 (0.50)	0.51 (0.50)	0.44
Female	0.60 (0.49)	0.63 (0.48)	0.47
Public insurance	0.62 (0.49)	0.65 (0.48)	0.37
Days with internet per week (1-7)	5.47 (2.71)	5.23 (2.84)	0.23
Days with social media per week (1-7)	5.37 (2.79)	5.19 (2.91)	0.37
Employed	0.62 (0.49)	0.64 (0.48)	0.53
Supports incumbent	0.50 (0.50)	0.51 (0.50)	0.86
Voted in previous election	0.73 (0.44)	0.74 (0.44)	0.68
Knows public pharmacy	0.61 (0.49)	0.67 (0.47)	0.09
Perceived relative price of public pharmacy	0.46 (0.18)	0.46 (0.23)	0.96
Perceived days to delivery at private pharmacy	8.80 (12.87)	8.35 (11.87)	0.61
Observations	319	507	

*Notes:* Columns (1) and (2) display the mean and standard deviation of different covariates at baseline for each experimental group. Column (3) displays the  $p$ -value from a test of equality of means across the groups.

TABLE A.10: EXPERIMENTAL RESULTS FOR ECONOMIC OUTCOMES

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A - Knowledge about public pharmacies									
	1(Knows about pharmacy)			log(Perceived price)			log(Perceived waiting time)		
Treatment	0.099*** (0.034)	0.069*** (0.026)		-0.117** (0.046)	-0.094** (0.045)		0.173 (0.107)	0.188* (0.103)	
Treatment $\times$ chronic			0.032 (0.033)			-0.114* (0.061)			0.134 (0.140)
Treatment $\times$ non-chronic			0.126*** (0.042)			-0.063 (0.065)			0.264* (0.151)
Dependent variable at baseline		0.489*** (0.039)	0.488*** (0.039)		0.382*** (0.049)	0.382*** (0.049)		0.397*** (0.068)	0.399*** (0.068)
Lee bounds	[-0.018, 0.134***]			[-0.236***, -0.020]			[0.049, 0.189]		
$p$ -value for $H_0: \beta_C = \beta_{NC}$	-	-	0.080	-	-	0.570	-	-	0.531
Mean for control group	0.773	0.773	0.773	9.070	9.070	9.070	1.387	1.387	1.387
Observations	514	514	514	498	491	491	445	425	425
R-squared	0.017	0.474	0.477	0.012	0.197	0.197	0.006	0.181	0.182
Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
County FE	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Panel B - Usage of public pharmacies									
	1(Enrolled)			1(Purchased)			Probability of usage		
Treatment	0.018 (0.024)	0.020 (0.024)		0.019 (0.017)	0.023 (0.018)		0.060* (0.035)	0.054 (0.036)	
Treatment $\times$ chronic ( $\beta_C$ )			0.032 (0.033)			0.043* (0.024)			0.085* (0.046)
Treatment $\times$ non-chronic ( $\beta_{NC}$ )			0.002 (0.034)			-0.008 (0.026)			-0.008 (0.057)
Knows pharmacy at baseline		0.050** (0.021)	0.050** (0.021)		0.015 (0.017)	0.015 (0.017)		-0.042 (0.043)	-0.045 (0.043)
Lee bounds	[0.007, 0.087***]			[0.015, 0.047***]			[0.060, 0.083]		
$p$ -value for $H_0: \beta_C = \beta_{NC}$	-	-	0.524	-	-	0.155	-	-	0.213
Mean for control group	0.069	0.069	0.069	0.028	0.028	0.028	0.540	0.540	0.540
Observations	514	514	514	514	514	514	387	387	387
R-squared	0.001	0.021	0.100	0.002	0.008	0.067	0.008	0.008	0.057
Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
County FE	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes

Notes: This table displays cross-sectional estimates using data from the field experiment. In particular, we present results using self-reported indicators about awareness and usage as dependent variables, on the treatment indicator and interactions with and indicator for chronic conditions. Columns 1, 4, and 7 include only a treatment indicator on the right-hand side; columns 2, 5, and 8 include the baseline level of the dependent variable, additional control variables, and county fixed effects; and columns 3, 6, and 9 add an interaction of the treatment indicator with an indicator for whether a member of the consumer household has a chronic condition. The set of control variables includes age and indicators for chronic condition, having completed high school education, female, and public insurance. Outcomes in Panel B either do not have baseline counterparts (which is the case by design of indicators for enrollment and purchase) or were not collected at baseline (which is the case for the probability of usage), so we instead control for knowledge of the public pharmacy at baseline. Reported Lee bounds are computed using only the treatment indicator as a covariate. Robust standard errors in parentheses. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

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