# Trade Associations and Collusion among Many Agents: Evidence from Physicians[††]

Jorge Alé-Chilet [*] Juan Pablo Atal[†]

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**Abstract**

We study a recent collusion case among physicians in Chile. Most gynecologists in one city formed a trade association to bargain for better rates with insurance companies. After unsuccessful negotiations, the physicians jointly terminated their insurer contracts and set a minimum price. We find that subsequent realized prices coincided with Nash-Bertrand prices, and that the minimum price was barely binding. We show that these actions ensured the association's stability and increased profits. Our findings show how a trade association may help sustain collusion among a large number of heterogeneous agents.

**Keywords**: Collusion, Trade associations, Minimum price, Physicians.

**JEL classification**: I11, L13, L41

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[*]Bar-Ilan University, Ramat Gan 5290002, Israel. E-mail: jorge.ale-chilet@biu.ac.il

[†]University of Pennsylvania, Department of Economics, USA. Axilrod Faculty Fellow. E-mail: ataljp@econ.upenn.edu

# 1 Introduction

Trade associations are prevalent in many professions, including engineering, law, and medicine. The role of these associations as information exchanges and standard setters has been largely acknowledged. However, trade associations also raise antitrust concerns as they may facilitate coordination on prices, establish barriers to entry, or undertake other activities that diminish competition (FTC, 2018, Kühn *et al.*, 2001). In fact, cases of successful collusion with a large number of agents typically involve the presence of a trade association, especially in differentiated product industries (Symeonidis, 2002; Levenstein and Suslow, 2006).

In this paper we study the strategies of a large number of agents that were able to coordinate successfully on consumer prices and on vertical negotiations through a trade association. Using detailed data on prices, sales, and court documents we provide the first empirical characterization of a trade association's collusive strategies. More explicitly, we empirically study the incentive compatibility of the Association and we provide empirical support for the theoretical predictions of collusion among heterogeneous firms (Harrington, 2016).

We analyze a trade association of 25 gynecologists that operated in Chile for 28 months before it was challenged by the National Economic Prosecutor and was ultimately abolished by the Supreme Court for collusive practices. The Association was created in mid-2011 and comprised 90% of the local gynecologists in one city. The Association's stated goal was to balance the bargaining power of its members vis-à-vis the insurance companies to achieve higher negotiated fees. Yet, the insurers were not willing to compromise. In response, six months after its creation, the members of the Association simultaneously terminated their contracts with the insurers —and therefore became out-of-network providers— and agreed to set their fees above a specified minimum price. As a result of these two measures, visits list and out-of-pocket prices rose, on average, by 80% and 200% respectively.

To rationalize the Association's pricing strategy and assess its stability, we first estimate patients' price elasticities for visits using the large increase in their out-of-pocket prices brought about by the Association. We use these demand elasticities to generate predictions for the visits prices from different supply-side models of associated doctors, including Nash-Bertrand competition, partial coordination, and full coordination.

Our main finding is that the realized prices of the physicians in the Association coincide with Nash-Bertrand prices. Moreover, we find that the minimum price was barely binding. Therefore, the Association's activities changed the insurer-provider network structure, but did not result in *supra-competitive* price levels.

We use our supply-side model to empirically investigate the incentive compatibility of the Association's collusive strategy. We estimate the counterfactual profits that each associated physician would have made by deviating from the Association. We find that these unilateral deviation profits are negative almost for every associated physician. Thus, the collusive strategy not only prevented deviations in prices while out-of-network, but it also prevented deviations with respect to the decision to leave the insurers' network. Therefore, Nash-Bertrand pricing in the out-of-network phase explains the stability of the Association.

Our findings provide empirical support for the theoretical insights of collusion among heterogeneous firms. First, the fact that the Association set a minimum rate instead of a unique fee is consistent with Harrington's 2016 insight that minimum prices preserve the heterogeneity in differentiated-product industries so that there always exists an incentive compatible (IC) minimum price at which firms can collude. Moreover, our empirical analysis shows that the Association's minimum price was roughly the static Nash-Bertrand of the lowest-priced firms, which Harrington (2016) shows to be an IC minimum price.

Our analysis provides several insights for antitrust practice. First, standard economic models and conventional wisdom predict that price coordination is hard with a large and heterogeneous set of firms (e.g., Motta, 2004; Levenstein and Suslow, 2006). This difficulty and the fact that physicians left the network and increased their fees only after extensive coordination suggest that communication through the trade association was essential to implement the joint contract termination. Thus, the case studied in this paper highlights the role of communication in cases of successful coordination.

Second, our results relate to the differences between economic collusion and illegal antitrust practices. Our finding of competitive prices implies absence of price collusion in the economic sense, while fixing a minimum price constituted an illegal practice *per-se* according to antitrust law (see Whinston, 2008, for a discussion).

Finally, our results have implications for the antitrust analysis of vertical relations. The negative welfare consequences for consumers brought about by the Association stemmed from the changes in the vertical relationship. Our estimates show that bargaining with insurance significantly restrained physician prices in the pre-association period. Thus, the Association harmed consumer welfare by appropriating the surplus that was captured by insurers and, hence, their clients.

**Related Literature** This paper builds on several strands of the literature. We contribute to the empirical literature on explicit collusion, such as Porter (1983), Levenstein (1997), Genesove and Mullin (1998), Genesove and Mullin (2001), Röller and Steen (2006), Asker (2010), Clark and Houde (2013), Igami and Sugaya (2017), and

Alé Chilet (2017).[1] We study a case of collusion among a large number of agents that involved the presence of a trade association. This is a common feature of collusion cases (Symeonidis, 2002; Levenstein and Suslow, 2006). To our knowledge, ours is the first paper that studies empirically the coordinating actions of a trade organization, and among them, the use of a minimum price.[2]

Moreover, we analyze collusion on dimensions other than price, namely the organization of the vertical network relationship. We find that non-price collusion was more important than price coordination in increasing profits. Similarly, Sullivan (2017) reports collusion on product characteristics.

Our methodology to study collusive equilibria after the breakdown of negotiations with insurers is due to Bresnahan (1987), Nevo (2001), Ciliberto and Williams (2014), and Miller and Weinberg (2017). We contrast the predictions of different assumptions on the nature of competition with the prices observed during the association period. We find that predicted Nash-Bertrand prices fit the Association prices well after the breakdown of the negotiations with the insurers.

This paper is also related to the literature on the anticompetitive effects of trade associations.[3] Donovan (1926) and Oliphant (1926), among others, discuss whether trade associations should be subject to antitrust law as a result of several antitrust suits in the 1920s. See also McGahan (1995) and Carnevali (2011). Symeonidis (2002) documents minimum price agreements in various British manufacturers as-

[1]Athey and Bagwell (2001), Athey and Bagwell (2008), and Athey *et al.* (2004) study collusion when prices are observed and firms receive private cost shocks. Harrington and Skrzypacz (2011) characterize an equilibrium where prices and quantities are private information, but firms truthfully report them to a trade organization.

[2]The use of a minimum price is a recurrent but understudied collusive strategy employed by trade associations of physicians and other professionals. Harrington (2016) reports collusion on minimum prices through trade associations for retail travel agents (Bingaman, 1996), specialty physicians (North Texas Specialty Physicians v. Federal Trade Commission, No. 06-60023, 2008 WL 2043040 (5th Cir., May 14, 2008)), and bus operators (Competition Commission of Singapore, 2009).

[3]Vives (1990) and Kirby (1988) discuss the role of trade associations as information exchanges.

sociations before antitrust legislation.[4] Levenstein and Suslow (2006) report that collusion among a large number of agents in unconcentrated industries usually arise due to the role of a trade association. Levenstein and Suslow (2011) argue that cartels involving trade associations are more visible, which makes them more likely to attract the scrutiny of antitrust agencies. Yet, at the same time, Levenstein and Suslow find that such cartels are also more stable. Our findings shed light on the reasons for the ob/gyn cartel stability.

Our work is also related to the literature on vertical relations, and in particular to insurer-provider negotiations in health care markets (Gowrisankaran *et al.*, 2015; Ho and Lee, 2017). Instead of studying agreement outcomes, we document a case where no agreement was reached and study the providers' pricing behavior in this event. Understanding such breakdowns is important because they shed light on the outside options of each side.[5]

# 2 Institutional Details

## 2.1 The Health Insurance Market in Chile

The Chilean health-care system is divided into a public and a private system. The focus of this paper on is the private system, which is a regulated health insurance market operated by a group of private insurance companies known collectively as Isapres ("Instituciones de Salud Previsional"). Isapres cover around 17 percent of the population. The public regime, FONASA, is a pay-as-you-go system financed

[4]Symeonidis also mentions that associations allowed individual price setting, especially in differentiated product industries, but maximum discounts to distributors were common.

[5]See Lee and Fong (2013). A strand of the literature in labor economics studies strikes and labor disputes. Card (1990), Gu and Kuhn (1998), and Cramton and Tracy (2003) provide theoretical models and empirical evidence. Classic references on collective bargaining and strikes are Hicks (1932) and Ashenfelter and Johnson (1969). Kennan (1986) and Cramton and Tracy (2003) review this literature.

by monthly contributions deducted from labor income, cost-sharing, and resources from the general government. FONASA covers roughly two thirds of the population (about 11 million people). [6]

Plans in the private system have two main coverage features: coinsurance rates (one for inpatient care and another for outpatient care) and coverage caps (insurer payment caps). Every plan assigns the insured a per-service payment cap, and these caps apply to each visit. Coinsurance rates and the insurer payment caps remain constant across visits and do not accumulate over time. For any particular claim, a person pays her coinsurance rate until the amount that the insurance company contributes reaches the cap for that service.

Individuals have access to different types of plans with respect to the provider network. "Preferred-provider" plans are tied to a specific network, although enrollees can use providers outside of their network at a higher price (similar to PPO in the US). Individuals can also choose—at a higher premium—plans with an unrestricted network of providers. Under these "free choice" plans, coverage is not tied to the use of a particular clinic or health care system, similar to a traditional fee for service indemnity plan in the United States. Companies also offer a small share of "closed network" plans, where enrollees can only use the services of the plan providers or must pay full price (the equivalent of the U.S. HMO).

Health care provision is also divided between public and private providers. In-network private providers negotiate their rates with insurance companies through bilateral negotiations, while out-of-network private providers set their rates privately.[7] On the other hand, FONASA reimburses private providers following a

[6] A small fraction of the population is insured by seven "closed" private insurance companies, which are available only to workers in certain industries; by special health care systems such as those of the Armed Forces, or do not have any coverage at all (Bitran *et al.*, 2010).

[7] With the exception of providers vertically integrated with insurance companies.

pre-determined rate for each procedure and a provider quality rank. [8]

## 2.2 The Antitrust Case

Gynecologists in Chillán, a city of 175,000 inhabitants in the Ñuble province in southern Chile, formed the Gynecologists' Association of Ñuble, legally constituted in August 2011.[9] The Association was formed by 25 out of the 29 Chillán's ob/gyn.

Upon the creation of the Association, all gynecologists in Ñuble held network contracts with Isapres that were negotiated independently. To balance the bargaining power of its members vis-à-vis the insurance companies, members of the association started to jointly negotiate their fees with Isapres in 2011. According to the documents presented in the antitrust case by the National Economic Prosecutor (Fiscalía Nacional Económica, FNE), in April 2011 the head of the Association, Dr. B., started approaching different Isapres and "informed them about the need" of increasing visit rates to CLP 41,000, up from an average price of CLP 14,000 (FNE, 2014 pp. 45-46).[10] After unsuccessful negotiations with the insurers, Dr. B. called the Association to meet for the first time in November 2011. The members agreed on a three-pronged plan: (1) canceling the members' individual contracts with the Isapres; (2) setting a minimum fee of CLP 25,000 (roughly USD 50) for visits; and (3) naming Dr. B. as the representative in future negotiations with the Isapres. Subsequently, all members of the Association sent termination letters to the Isapres, which would go into effect in January 21, 2012. The Isapres faced customers com-

[8]The provider quality rank goes from 1 to 3 to adjust for quality and complexity differences.

[9] Chile is divided administratively into 15 regions, each subdivided into 54 provinces. The southern region "Biobío" is divided into four provinces: "Concepcion" is the largest province in the Region, followed by "Ñuble", "Biobío" and "Arauco".

[10]During our period of study, the exchange rate was roughly CLP 500 = USD 1.

plaints, but did not accept the physicians' requests immediately.[11] Throughout 2012 the insurance companies contacted the physicians separately and offered rates lower than those requested by the Association. The physicians rejected those private offers.[12] It was not until March 2013 that a major Isapre accepted (almost fully) the Association's terms.[13] Yet, that same month the FNE started investigating the Association for antitrust offenses and filed an indictment in October 2013. The Competition Tribunal found the Association members guilty of colluding in 2015. Finally, the Supreme Court ordered the dissolution of the Gynecologists' Association in 2016.

Figure 1 shows the evolution over time of the list price (before reimbursement) and coinsurance rate for the average gynecologist-insurer pair in Chillán as well as in two of its main neighboring cities. Panel (*a*) shows that the list price in Chillán increased by 80 percent during the Association period, from roughly CLP 12,900 to CLP 23,300. The termination meant not only that the physicians rose their rates, but also that these rates would be covered as out-of-network visits. Panel (*b*) shows an average increase of 65 percent on the average coinsurance rate, from 26 percent to 43 percent. Combined, the out-of-pocket cost of a visit to the average doctor in Chillán increased by 200%.[14] There were no discernible changes in the out-of-pocket costs in neighboring cities.

Compliance within the Association with the minimum visit price was substan-

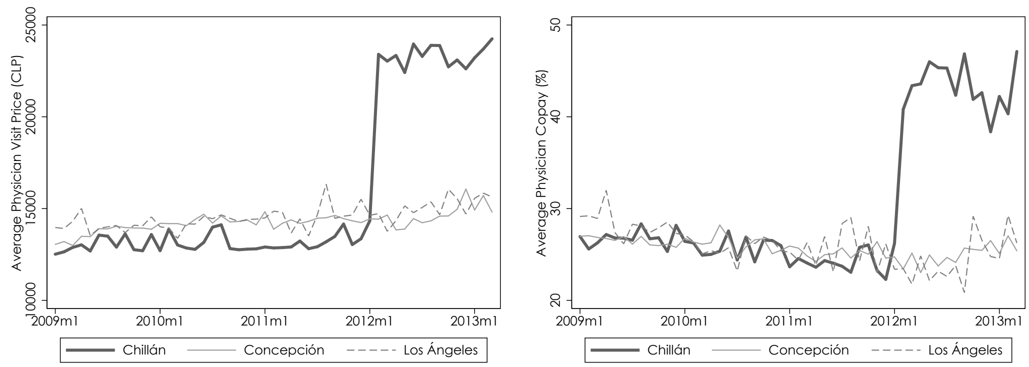
[11] The agreement also included a minimum price for surgeries of 4-4.4 times the FONASA rates. Yet, the agreement was not very effective for surgeries. For this reason, and because surgeries require the presence of a larger medical team, we focus on visits.

[12] FNE (2014) pp. 68-69. Foreseeing these proposals, Dr. B. reminded the Association members that "any information and/or negotiation with the Isapres should be undertaken through its president" (FNE, 2014 p. 55).

[13] The agreed terms were CLP 25,000 for visits and 4 times the rate paid by FONASA for surgeries. In addition, a small Isapre closed to public enrollment, reached an agreement with the association in January 2012 of CLP 20,000 for visits and 4.4 times the FONASA rate for surgeries.

[14] As we discuss in the next section, the list price and coinsurance increases were offset by a shift to physicians that did not raise fees, so that the mean visit out-of-pocket price rose by 60 percent. Patients were surprised of the change, which was reported in the local media (FNE, 2013, p.9)

**Figure 1:** Prices and Coinsurance rates over time




**(a)** Average Visit Price **(b)** Average Coinsurance Rate

Note: The Figure shows the median visit price and coinsurance rate faced by patients of the average gynecologists in Chillán and in two other major neighboring cities (Concepción and Los Ángeles). Prices are in Chilean Pesos (approximately CLP 500 = USD 1).

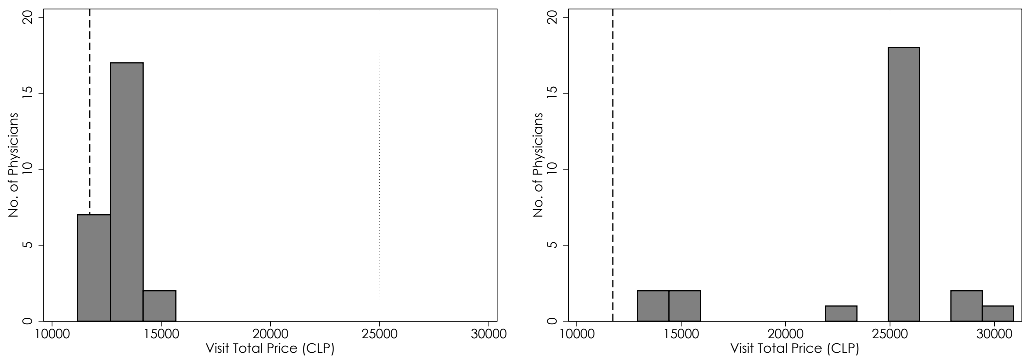
tial. Figure 2 shows a histogram of each physician's median list price in 2011 and 2012. The dashed line shows the FONASA rate and the dotted line indicates the minimum price set by the Association. The price distribution shifts right in 2012, to the extent that the mode price in 2012 was almost twice as high as the mode price in 2011. That was the case even for the most expensive doctors, who more than doubled their list price. In addition, as we show in Section 4, those physicians who did not raise their rates greatly increased their market share.[15]

# 3 Data

We use the two main data sources of the antitrust case. The first source is the insurance companies' administrative data. It includes visits and surgeries that were registered by the Isapres between January 2009 and March 2013 for gynecologists

[15]Anecdotally, one of the physicians who did not join the Association was the former government's director of health services in Ñuble and, according to our data, started working in Chillán's private sector only in mid-2010.

**Figure 2:** Histogram of Median Prices




**(a)** Pre-Association **(b)** Association

Note: The Figure shows the median price distribution of Chillán gynecologists in the period before and after collusion (2011 and 2012). We drop prices in the transition month of January 2012. Prices are in Chilean Pesos (approximately CLP 500 = USD 1). The dashed line is the FONASA rate and the dotted line is the set minimum price.

operating in Chillán and its neighboring provinces. Each record contains physician and medical establishment identifiers, a scrambled patient identifier, patient's province of residence, date, price, out-of-pocket expenditures, and a procedural code. The second dataset contains the receipts issued by the indicted physicians for the period 2012. It includes patient and physician identifiers and the total amount payed. We use receipts information because many patients did not process their reimbursement with the insurance companies after doctors became out-of-network and hence their visits are not included in the first dataset.[16] Since we estimate aggregate demand at the insurer-physician-month triad, we calculate the price and coinsurance rate as their average over the visits corresponding to each combination. Finally, we obtain the physicians' main address from the case documents

[16]We cannot match the two datasets at the patient level. For the demand estimation, we assume that these visits were reimbursed at the out-of-network coverage rate. We also have receipts data for 2011, which we do not use in order to avoid having duplicates as a result of pooling the two datasets. In addition, we drop the physician-month observations in 2012 that are likely to be the result of incorrect imputation. For example, in the case of one physician two months register 35% of all yearly receipts.

(FNE, 2013).[17]

Figure 3 shows the evolution of prices and profits over time for associated and non-associated physicians. As as a measure of a physician's marginal cost we use the FONASA rate.[18] The figure shows that concurrently with the price increase among associated physicians there was a large decrease in their average number of visits, as well as a large increase in the average visits among non-associated doctors. In Appendix A3, we show evidence that patient's switching rates across physicians increased on average by 32 percent after the collusion. Moreover, switching occurred mostly from colluding towards non-colluding physicians.

Pre-collusion profits of associated and non-associated doctors are stable, and similar across groups. However, profits change drastically for both groups after the association formation. We find that collusive profits were on average 4 to 5 times larger than in the pre-collusive period, both for associated and non-associated doctors.

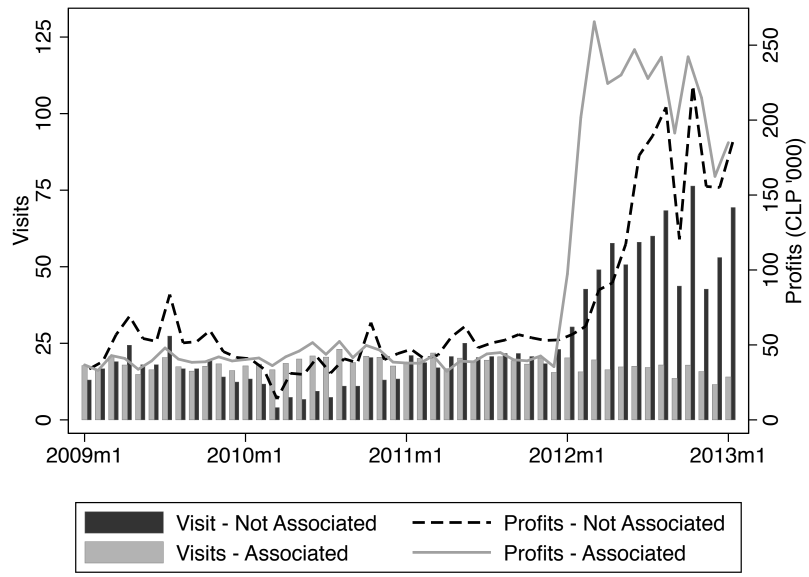
# 4 Demand

## 4.1 Demand Model

Our empirical strategy to estimate demand leverages the price shock brought about by the gynecologists' joint contract termination, which subsequently increased list prices and coinsurance rates. We model physicians as differentiated-product firms to allow for idiosyncratic preferences for ob/gyns, which is likely an important feature of the industry. We assume a nested logit demand model, where the set of

[17]For the physicians who were not indicted, we use local websites.
[18]We discuss this assumption in more detail in Section 5.1

**Figure 3:** Average Number of Visits and Profits over Time




Note: The Figure shows the average number of visits and profits (using Fonasa as the cost) for associated and not associated physicians.

physicians is partitioned into non-overlapping nests $B_k$, $k = 1, \dots, K$.[19]

The utility of patient $p$ enrolled in insurer $j$ for visiting doctor $i$ in nest $k$ is

$$u_{pijt} = \delta_{ijt} + \epsilon_{ijt} + \eta_{pkt} + (1 - \sigma)\nu_{pijt},$$

where $\delta_{ijt}$ denotes the mean utility of insurees who visited doctor $i$ at time $t$, $\epsilon_{ijt}$ is the unobserved mean-utility component of physician desirability, $\eta_{pkt}$ is an unobserved common shock to physicians in nest $k$, $\sigma$ is a parameter that measures within nest correlation, and $\nu_{pijt}$ is an idiosyncratic shock. The model assumes that $\eta_{pkt} + (1 - \sigma)\nu_{pijt}$ has a generalized extreme value distribution.

The share of insurees who visit doctor $i$ in time $t$ (among all female insurees in

[19]In practice, these nests are constructed based on the doctor's geographic location as explained in detail in Section 4.2. The goal of these nests is to capture differentiation across doctors arising from patient's heterogeneous preferences for location.

insurance company $j$), follows the standard share equation:

$$s_{ijt} = \frac{\exp(\frac{\delta_{ijt}}{1-\sigma}) \sum_{h \in B_k} \left[\exp(\frac{\delta_{hjt}}{1-\sigma})\right]^{-\sigma}}{1 + \sum_{l=1}^{K} \sum_{h \in B_l} \left[\exp(\frac{\delta_{hjt}}{1-\sigma})\right]^{1-\sigma}}. \tag{1}$$

We parameterize $\delta_{ijt}$ as

$$\delta_{ijt} = \alpha p_{ijt}(1 - c_{ijt}) + \mu_{ij} + f(t), \tag{2}$$

where $p_{ijt}$ is the list price of a visit to physician $i$ of enrollees of $j$ in period $t$, $1 - c_{ijt}$ is the coinsurance rate, $\mu_{ij}$ represent fixed effects for doctor-insurer pairs; and $f(t)$ is a flexible function of time that captures time-specific common shocks to demand.

Let $s_0$ represent the share of the outside option, and $s_{ijt/B_k} = s_{ijt} / \sum_{h \in B_k} s_{iht}$, $j \in B_k$, be the *within*-nest share of physician $i$. Then we can invert equation (1) as in Berry (1994) to obtain

$$\ln(s_{ijt}) - \ln(s_{0jt}) = \alpha p_{ijt}(1 - c_{ijt}) + \mu_{ij} + f(t) + \sigma \ln(s_{ijt/B_k}) + \epsilon_{ijt} \tag{3}$$

with our main objects of interest being the parameters $\alpha$ and $\sigma$.[20]

## 4.2 Estimation

The nested logit model requires grouping physicians into nests according to predetermined susbstitutability patterns. We assume that patients are more likely to substitute to a physician that is either in the same medical center or has a practice

[20] We assume that physicians capacity constraints do not bind. Binding capacity constraints in the data would imply that we underestimate the price coefficient. Moreover, in the supply model binding capacity constraints reduce both potential deviation profits (Staiger and Wolak, 1992) and the severity of punishment (Brock and Scheinkman, 1985; Fabra, 2006). We provide evidence of absence of capacity constraints, both in the data and in the predictions of the model in Appendix A5.

in a nearby location (see e.g. Phelps and Newhouse, 1974). Hence, we construct six nests based on the location of each doctor's practice.

Figure A1 of the Appendix shows a map with the location of physicians and the resulting nests. There are three large medical centers where five or more doctors co-locate. We assume that each of those centers constitute a different nest. We group the rest of physicians –who either are a single practice or are co-located with at most two others– into three other nests based on geographic distance. Also, we construct two additional nests, one for physicians outside Chillán, and another one for the outside option.

The estimation of equation (3) poses two challenges. First, 26 percent of market share observations in our data are equal to zero. This fact imposes a type of censoring in the estimation of equation (3) because the logit model does not allow for zero shares. Second, we face the standard endogeneity problem of prices and within shares, as both $\epsilon_{ijt}$ and the within shares are correlated with prices. We address both issues below.

**Zero shares** As noted in Section (2), all doctors are in all insurees' choice set. Insurees can in principle see any given doctor by paying the corresponding out-of-pocket price. Therefore, the observed zero shares $s_{ijt}$ can be interpreted as the realization of small shares in a *small-sample*, which differ from the true population probabilities implied by the model. To solve this issue we rely on Gandhi *et al.* (2017), who propose an asymptotic correction based on the assumption that sales distribute according to Zipf's Law. This implies that for every insurer $j$

$$M_j \frac{s_{ijt}}{1 - s_{0jt}} | N, M_j, s_{0jt} \sim DCM \left( \theta_j 1_N, M_j (1 - s_{0jt}) \right), \tag{4}$$

where $M_j$ is the market size, $N$ represents the number of doctors and $DCM(\cdot)$, is the Dirichlet Multinomial Distribution with the same $N$ parameters $\theta_j$. Denoting the digamma function by $\psi(\cdot)$, we estimate the following version of the nested logit demand model

$$\begin{aligned}\psi(\theta_j + M_j s_{ijt}) - \psi(M_j s_{0jt}) = \alpha p_{ijt}(1 - c_{ijt}) + \mu_{ij} + f(t) + \\ \sigma \left[ \psi(\theta_j + M_j s_{hjt}) - \psi(\sum_{h \in k} \theta_j + M_j s_{ijt}) \right] + \epsilon_{ijt}.\end{aligned} \tag{5}$$

In this equation, the left hand side correspond to the expectation of the logit dependent variable of Equation (3), and the term in square brackets is the expectation of the log within share. The $DCM$ parameters $\theta_j$ can be estimated from (4) via Maximum Likelihood since $N$, $s_{0jt}$, and $M_j$ are known.[21]

**Endogeneity of Prices and Shares** The estimation of Equation (5) by ordinary least squares (OLS) will result in biased estimates in the presence of correlation between the explanatory variables and the unobserved demand shocks $\epsilon_{ijt}$. We identify the own- and cross-price demand elasticities using the large increase in out-of-pocket expenditures that stemmed from the gynecologists' contract termination. In particular, we assume that the emergence and the membership in the Association was not a result of idiosyncratic demand shocks, after controlling for the fixed effects.

Two pieces of evidence support this assumption. First, as shown in Figure 1, prices and coinsurance rates in Chillán had a similar level and followed a similar trend to those in neighbouring cities before the Association emerged in Chillán. Thus, the data suggests that the emergence of the Association in Chillán was not a

[21] In practice, we set the market size equal to the number of female insurees in each Isapre.

result of particular conditions in this market. Moreover, Figure 3 shows that profits and visits followed a similar trend in the pre-Association period for associated and non-associated doctors. Therefore, the data shows that the membership in the Association does not seem to have been driven by doctor-specific shocks to demand.

Under this assumption, the Association period serves as an instrumental variable (IV) that we use to solve the endogeneity problem. In particular, we use as instruments an indicator for the post-agreement period (after February 2012) and its interaction with a dummy variable for whether physician $i$ joined the agreement. Hence, we identify the elasticities using changes in prices and coinsurance rates within the doctor-insurer pair before and after the price change. [22]

### 4.3 Results

Table 1 presents the results of estimating demand. All specifications use a month-of-the-year fixed effect to control for seasonality and year fixed effects to allow for common trends. We present first the OLS and IV results in Columns (1) and (2) of a simpler logit model estimated from Equation (5) under the constraint $\sigma = 0$. Columns (3) and (4) present the OLS and IV nested logit estimation results of Equation (5).[23]

In our main specification (Column (4) of Table 1) the estimates of $\alpha$ and $\sigma$ are significantly different from zero. Also, $\sigma \in [0, 1]$, which is consistent with utility maximization.[24] The Anderson-Rubin and Angrist and Pischke (2008) $F$-tests reject that the instruments are weak. The rows under the estimates show the average own

[22]Our instruments use a supply-side event to identify the demand as in Porter (1983), Eizenberg and Salvo (2015), and Alé Chilet (2017).

[23]Appendix A2 presents additional demand estimates resulting from standard ways of correcting for zero market shares, like dropping observations with no sales or replacing the zeros with an arbitrary positive number. See Gandhi *et al.* (2017) for a discussion on the drawbacks of those methods.

[24]Given that the association period, which we use as an instrument, has significant overlap with the 2012 year fixed effect, we also tried fixed effects for 8 month periods. The results are unchanged.

elasticity in the period before and after the Association was formed.[25]

**Table 1:** Demand Estimates

| | (1) Logit OLS | (2) Logit IV | (3) NL OLS | (4) NL IV |
|---|---|---|---|---|
| $p_{ijt}(1-c_{ijt})$ | 0.004 | -0.146*** | 0.004 | -0.050*** |
| | (0.010) | (0.017) | (0.004) | (0.018) |
| $\ln within - share$ | | | 1.040*** | 0.735*** |
| | | | (0.005) | (0.152) |
| $\eta_{pre}$ | 0.02 | -0.60 | -0.29 | -0.65 |
| $\eta_{post}$ | 0.04 | -1.53 | -0.75 | -1.68 |
| Observations | 6620 | 6620 | 6620 | 6620 |
| AR $F$-stat | | 39.42 | | 39.42 |
| AP $F$-stat 1 | | 1899.99 | | 334.84 |
| AP $F$-stat 2 | | | | 19.83 |

Note: The Table shows the demand estimates. All specifications include month-of-the-year and year fixed effects. Out of pocket expenditures are in thousand CLP. The AR $F$-stat corresponds to the Anderson-Rubin Wald $F$-statistic. The AP $F$-stat 1 and 2 correspond to the first-stage $F$ statistics of the excluded instruments for out-of-pocket expenditure and within shares respectively (Angrist and Pischke, 2008). Heteroscedasticity robust standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$

# 5 The Association's Pricing Strategy

## 5.1 Model

We model physicians during the Association period as differentiated-product firms, which may or may not collude with their competitors.[26] This allows us to calculate competitive and collusive prices given the demand estimates.

[25] Note that the own elasticity for the logit and nested logit models are quite similar; yet, as expected, there is a large difference in the cross elasticities (not reported).

[26] We do not model bargaining with the insurers during the agreement period because physicians terminated their contracts with the insurers. Moreover, the supply-side models we use are similar to a case in which physicians have all the bargaining power in a potential negotiation.

In every period each of the $N_c$ colluding physicians decides a unique list price for all insurance companies. Colluding physicians take as given the prices of the $N - N_c$ non-colluding physicians —which are set non-strategically at their pre-agreement level— and the out-of-network coinsurance rates for each insurer. We denote by $\mathbf{p}$ the vector with the stacked prices of all the colluding physicians, $\mathbf{c}_j$ the vector of coverage rates (1 minus coinsurance rates) in insurer $j$ and the corresponding vector of monthly visits $\mathbf{q}_j(\mathbf{p}, \mathbf{c}_j)$. The first-order condition of colluding physicians can be written in matrix form for each period as

$$\mathbf{p} = \mathbf{mc} - \left( \Omega^* * \sum_j E_j * C_j \right)^{-1} \sum_j \mathbf{q}_j(\mathbf{p}, \mathbf{c}_j), \tag{6}$$

where the operator $*$ denotes element-by-element multiplication.[27] The matrix $E_j$ corresponds to the own- and cross-derivatives of demand with respect to out-of-pocket prices in insurer $j$, which we estimated in Section 4. $C_j$ is a matrix of coinsurance rates with elements $C_j(i, s) = 1 - c_{ij}$, that converts the list prices into out-of-pocket prices for visits to doctor $i$ among insurees of Isapre $j$. Finally, the vector $\mathbf{mc}$ is the physicians' marginal cost. We assume that $\mathbf{mc}$ is constant across physicians, and is equal to the FONASA rate, which is determined yearly by the regulator. The rationale for this assumption is that the FONASA rate corresponds to the opportunity cost of seeing one extra privately insured patient.[28]

We also define the ownership matrix $\Omega^*$ for the colluding physicians. The ownership matrix specifies the degree to which each colluding physician internalizes

[27] We omit time indexes to simplify notation.

[28] The average marginal cost in the pre-association period is equal to CLP $10,970$. One could argue that a visit of a privately insured patient is more costly than a visit of a publicly insured patient if, for instance, doctors spend more time with the privately insured patients. In that case, the marginal cost would be bounded from below by the FONASA rate, and from above by the actual price, that is, CLP $10,970$ and CLP $13,982$, respectively. In the extreme case that the pre-agreement price was equal to the doctor's marginal cost, we would underestimate the equilibrium prices by CLP $3,000$, or 9 percent of the Nash Price.

other physicians' profits. As in Nevo (1998) we focus on a general ownership matrix to accommodate different degrees of coordination on prices among colluding physicians, parametrized with a scalar $\kappa \in [0, 1]$ such that

$$\Omega^* = \begin{bmatrix} 1 & \kappa & \dots & \kappa \\ \kappa & 1 & \dots & \kappa \\ \vdots & \vdots & \ddots & \vdots \\ \kappa & \kappa & \dots & 1 \end{bmatrix}. \tag{7}$$

When $\kappa = 0$, the matrix $\Omega^*$ becomes the identity matrix, which determines the Nash equilibrium; any $\kappa > 0$ determines a different partial collusive equilibrium, with a higher $\kappa$ indicating a higher degree of coordination on prices. The extreme of $\kappa = 1$ corresponds to full collusion on prices.[29] For any given assumed value of $\kappa$, we solve for $\mathbf{p}$ by numerically finding the fixed point in Equation (6).

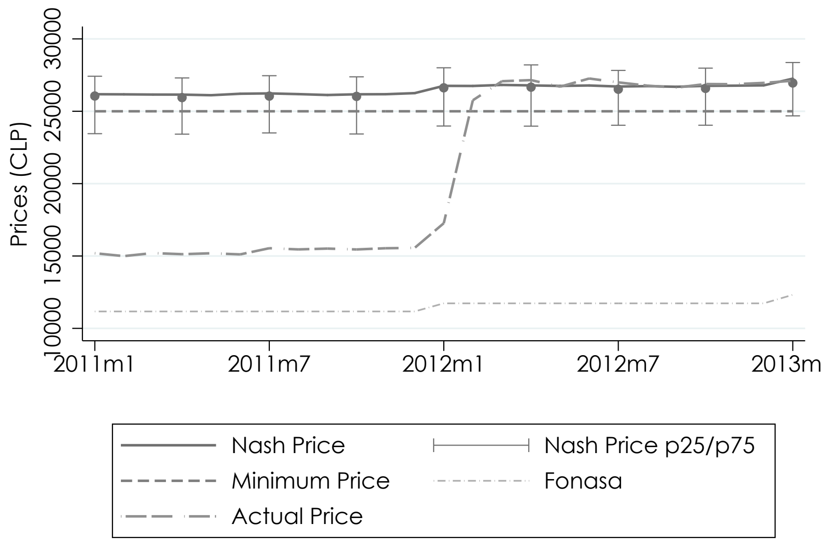
### 5.2 Results

The main results of the predicted price from Equation (6) are in Figure 4. The figure presents the time series of equilibrium prices assuming Nash-Bertrand competition among associated doctors ($\kappa = 0$).[30] We plot (quantity-weighted) average Nash prices, as well as the 25 and 75 percentiles of the distribution of prices across colluding physicians. For comparison, the figure also includes the average observed prices, the marginal cost (the FONASA rate), and the minimum price set by the Association (CLP 25,000). In addition, Figure 5 shows the distribution of average Nash prices of each physician during the Association period.

[29] We refer to Miller and Weinberg (2017) for the discussion of whether $\kappa$ can be interpreted as a conduct parameter. See the references there, especially Corts (1999) and Sullivan (2017).

[30] Since doctors left the network, we use out-of-network coinsurance rates, which we calculate as the average coinsurance rate of each insurer-physician pair across out-of-network visits in the Association period.

**Figure 4:** Simulated and Actual Prices

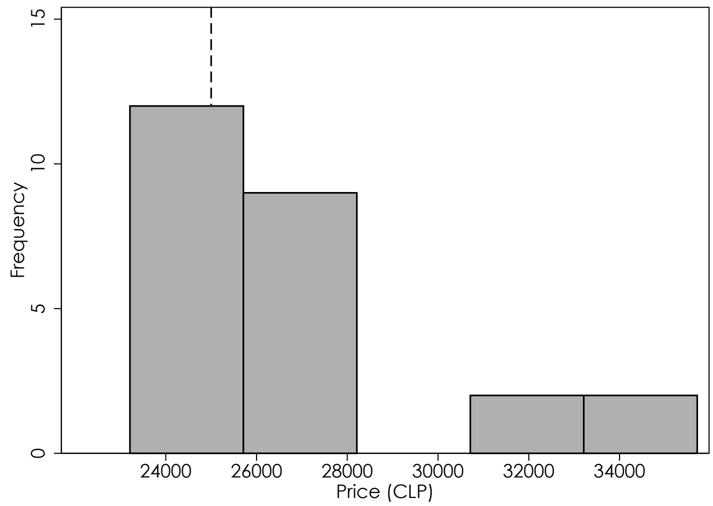



Note: The Figure shows the Nash prices for out-of-network coinsurance rates, the average actual price, and the FONASA rate. The dashed line shows the minimum price set by the Association. In addition, the vertical lines show the 25 and 75 percentiles of the out-of-network Nash prices, and the dot indicates the median.

Three facts about our results are particularly noteworthy. First, the average Nash-Bertrand price (CLP 26,800) is almost twice as high as the observed price in the pre-agreement period. This difference reflects the low bargaining power of the doctors vis-à-vis insurance companies in the pre-Association period. Second, as shown in Figure 4, the average Nash-Bertrand price is almost equal to average price set by the members of the Association. By contrast, the collusive price ($\kappa = 1$) is equal to CLP 52,336; twice as high as the average observed price. Third, as shown in Figure 5, the distribution of the Nash-Bertrand prices is such that the minimum price barely binds.

Our results provide two main insights. First, the Association served as a collusive device on the physicians' decision to become out-of-network providers, but not to set *supra-competitive* prices. Thus, the price increase during collusion stems

**Figure 5:** Distribution of Nash-Bertrand Prices




Note: The Figure plots the histogram of the estimated Nash prices of each associated physician, averaged over the Association period. The dashed line shows the Association's minimum price.

almost fully from the fact that the physicians were no longer constrained by their previous agreements with the insurers. Second, our findings provide empirical support for the existence of minimum prices as an effective collusive mechanism among heterogeneous firms. In particular, Harrington (2016) shows that a minimum price arbitrarily close to the static Nash-Bertrand price of the lowest-priced firm is IC. This is similar to what we find empirically: In our case, the physicians' lowest Nash price was CLP 23,210, just below the minimum price.

### 5.3 Stability of the Association

In this subsection we evaluate the stability of the Association by analyzing whether colluding physicians would have profited from leaving the Association unilaterally. In fact, insurers approached physicians individually during the Association period and offered them to rejoin their network. Yet, the Association prevailed and no physician deviated. Finding a large number of physicians with positive deviation profits would entail that the Association was not stable or depended critically on non-monetary incentives. On the contrary, finding negative deviation profits provides support to our model as a characterization of a stable collusive agreement.

We evaluate the decision to stay in the Association under the assumption that prices out-of-network are set in the Nash-Bertrand game, which is consistent with our findings in the previous section. We assume that deviating physicians set their average pre-Association price, that is, we assume they return to their original insurer contract. Formally, consider the profits for physician $i$,

$$\pi_i = \sum_j (p_i - mc)\, q_j(p_i, \mathbf{p_{-i}}, c_{i,j}, \mathbf{c}_{-i,j}),$$

where $p_i$ is the price of doctor $i$, and $\mathbf{p_{-i}}$ is the vector of prices of all doctors except for $i$, and $j$ indexes the insurance companies. In the Association agreement, prices are equal to the Nash-Bertrand prices, $\mathbf{p^N}$, and coverage rates are at the out-of-network level $c^O$. Therefore, the profits for physician $i$ in the Association are given by:

$$\pi_i^N = \sum_j \left(p_i^N - mc\right) q_j(p_i^N, \mathbf{p^N_{-i}}, c^O_{i,j}, \mathbf{c^O}_{-i,j}).$$

Consider doctor $i$'s deviation from the Association. The vector of prices when doctor $i$ deviates is $\left(p_i^d, \mathbf{p^N_{-i}}\right)$, where $\mathbf{p^N_{-i}}$ is the vector of Nash prices for all doctors except for $i$, and $p_i^d$ is the in-network negotiated price of doctor $i$. Moreover, patients of doctor $i$ receive in-network coverage, $c^I_{i,j}$, such that the extra profits from deviating are equal to

$$\pi_i^d - \pi_i^N = \sum_j \left(p_i^d - mc\right) q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^I, \mathbf{c^O_{-i,j}}) - \sum_j \left(p_i^N - mc\right) q(p_i^N, \mathbf{p^N_{-i}}, c_{i,j}^O, \mathbf{c^O_{-i,j}}). \tag{8}$$

When physician $i$ deviates, she lowers the price from $p_i^N$ to $p_i^d$, but also profits from a higher coverage rate for her patients. The two effects can be decomposed by re-writing equation (8) as

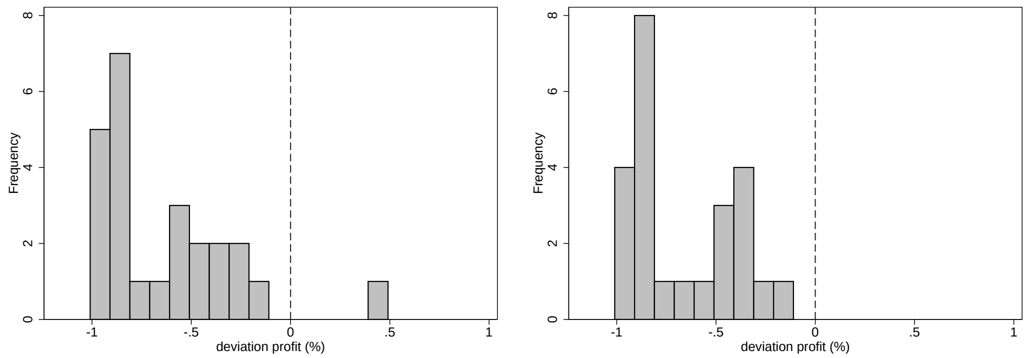
$$\pi_i^d - \pi_i^N = \underbrace{\sum_j \left(p_i^d - mc\right) q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^O, \mathbf{c^O_{-i,j}}) - \sum_j \left(p_i^N - mc\right) q_j(p_i^N, \mathbf{p^N_{-i}}, c_{i,j}^O, \mathbf{c^O_{-i,j}})}_{\text{Out-of-network price deviation from Nash Pricing} < 0} + \underbrace{\sum_j \left(p_i^d - mc\right) \left(q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^I, \mathbf{c^O_{-i,j}}) - q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^O, \mathbf{c^O_{-i,j}})\right)}_{\text{Higher sales from higher coverage} > 0} \tag{9}$$

The first term of equation (9) is the change in profits due to price deviation while keeping coverage rates at the out-of-network level. Since prices are set at the Nash-Bertrand level for out-of-network coverage, the first term in equation (9) is negative. On the other hand, the second term is the deviation towards being outside of the network. This term is positive, and proportional to the extra visits generated by higher coverage in-network.[31] Therefore, deviation profits have an ambiguous sign, so we quantify them empirically using our demand estimates.

Panel (a) of Figure 6 presents the results. The panel shows the histogram of deviation profits across physicians, calculated as a share of the collusive profits,

[31] It is easy to show that $q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^I, \mathbf{c^O_{-i}}) - q_j(p_i^d, \mathbf{p^N_{-i}}, c_{i,j}^O, \mathbf{c^O_{-i,j}}) \simeq -E_{ij} \times p_i^d \times \left(c_i^H - c_i^L\right) > 0$ where $E_{ij}$ is own-price elasticity of demand for doctor $i$ among enrollees of $j$ with respect to the out-of-pocket price.

**Figure 6:** Deviation Profits and Association Stability




**(a)** Step 1 **(b)** Step 2

The Figure shows the distribution of per-period deviation profits across physicians. Each point in the figure corresponds to the deviation profit of a different physician, as a percentage of the profit from staying in the Association.

$(\pi^N - \pi^d)/\pi^N$. Each unit of observation corresponds to a different deviating physician (and, therefore, to a different counterfactual scenario) among the 25 colluding physicians.

We find positive deviation profits for only one of the 25 colluding physicians. The median doctor would have lost 79 percent of her collusion profits by deviating to their pre-agreement price and staying in-network.

In order to assess the overall stability of the Association, we iterate this process by removing the doctor with positive deviation profits and recalculating the new equilibrium prices and profits from unilateral deviations until no doctors have positive deviation profits. This process stops in the second iteration, in the counterfactual situation where the doctor with positive deviations profits in Panel (a) leave the Association as it was actually formed.

The results of each iteration step are in Panel (b) Figure 6. Panel (b) shows the histogram of unilateral deviation profits after the doctor with positive deviation profits in Panel (a) deviates. In this case, none of the remaining doctors have incentives to deviate.

These findings provide evidence of the incentive compatibility of the Association's collusive strategy. This strategy not only prevented deviations in prices while out-of-network, but it also prevented deviations with respect to the decision to leave the insurers' network.

**Incentives to join the Association** An alternative interpretation of equation (9) is that it quantifies physician $i$'s incentives to join the Association, assuming that everyone else joins and charges prices $\mathbf{p}^N_{-i}$. However, it might be natural to assume that each physician considers her impact on the equilibrium premiums when considering whether to join the Association or not. In that case, each physician $i$ computes her profits from not joining, $\tilde{\pi}^d$, that depend on the outcome of the Nash-Bertrand pricing game when physician $i$ does not join, $\tilde{\mathbf{p}}^N_{-i}$. In Appendix A6 we show the resulting deviation profits after considering endogenous repricing of the associated physicians. We find that price adjustments are small so that deviation profits in this case are similar to the results of Figure 6. We interpret this finding as showing that joining the Association was IC.

## 6 Conclusion

In this paper we study the collusive strategies of a trade association of physicians. The Association members undertook two coordinating strategies: joint termination of their contracts with insurance companies and agreement on a minimum price per visit. These joint measures were effective in increasing the members' profits.

We find that the realized prices coincided with the competitive price and that the minimum price was barely binding. These findings suggest that the change in the structure of the physician-insurer relationship was more important to increase physicians' profits than price coordination. We also estimate unilateral deviation profits from joint contract termination, which we find to be much lower than the collusive profits. We conclude that the Association's collusive strategy was incentive compatible. These results conform with the theoretical predictions of collusive behavior among heterogeneous firms.

Our paper has antitrust policy implications. The Gynecologists' Association did not manage to raise prices much above the competitive outcome. However, the Association was highly successful in coordinating the bargaining efforts vis-à-vis the insurers. This result suggests that the nature of coordination in negotiations in vertical relationships is different from that of coordination in prices. While it is well known that price collusion among a large number of heterogeneous agents is difficult, the case studied in this paper suggests that such hurdles might not arise in other types of coordination. Antitrust authorities should provide sufficient scrutiny to such events.

Furthermore, the case studied in our paper highlights the role of communication through a trade association in cases of successful coordination. Although the presence of communication is central in antitrust practice, its role in collusion is not well understood by theoretical models.

Finally, the failed negotiation documented in this paper constitutes a departure from the agreement equilibria usually studied in the empirical industrial organization literature. We interpret our findings as showing that coordination led to increased bargaining power, which ultimately resulted in a breakdown of the physician-insurers negotiation. Documenting such breakdowns is important because they shed light on the outside options of each side, and provide theoretical support for the modeling of equilibrium bargaining outcomes. Understanding the dynamics of a failed bargaining process is an interesting area for future research.

## References


ALÉ CHILET, J. (2017). Gradually rebuilding a relationship: The emergence of collusion in retail pharmacies in Chile. *Working Paper*.

ANGRIST, J. D. and PISCHKE, J.-S. (2008). *Mostly harmless econometrics: An empiricist's companion*. Princeton University Press.

ASHENFELTER, O. and JOHNSON, G. (1969). Bargaining theory, trade unions, and industrial strike activity. *American Economic Review*.

ASKER, J. (2010). A study of the internal organization of a bidding cartel. *American Economic Review*, **100** (3), 724–62.

ATHEY, S. and BAGWELL, K. (2001). Optimal collusion with private information. *RAND Journal of Economics*, pp. 428–465.

— and — (2008). Collusion with persistent cost shocks. *Econometrica*, **76** (3), 493–540.

—, — and SANCHIRICO, C. (2004). Collusion and price rigidity. *The Review of Economic Studies*, **71** (2), 317–349.

BERRY, S. T. (1994). Estimating discrete-choice models of product differentiation. *The RAND Journal of Economics*, pp. 242–262.

BINGAMAN, A. (1996). Recent enforemcent actions by the antritrust division against trade associations. 32nd Annual Symposium of the Trade Association and Antitrust Law Committee of the Bar Association of the District of Columbia, US Department of Justice, Antitrust Division.

BITRAN, E., ESCOBAR, L. and GASSIBE, P. (2010). After Chile's health reform: Increase in coverage and access, decline in hospitalization and death rates. *Health Affairs*, **29** (12), 2161–2170.

BRESNAHAN, T. F. (1987). Competition and collusion in the American automobile industry: The 1955 price war. *The Journal of Industrial Economics*, pp. 457–482.

BROCK, W. A. and SCHEINKMAN, J. A. (1985). Price setting supergames with capacity constraints. *The Review of Economic Studies*, **52** (3), 371–382.

CARD, D. (1990). Strikes and wages: a test of an asymmetric information model. *The Quarterly Journal of Economics*, **105** (3), 625–659.

CARNEVALI, F. (2011). Social capital and trade associations in america, c. 1860–1914: a microhistory approach. *The Economic History Review*, **64** (3), 905–928.

CILIBERTO, F. and WILLIAMS, J. W. (2014). Does multimarket contact facilitate tacit collusion? Inference on conduct parameters in the airline industry. *The RAND Journal of Economics*, **45** (4), 764–791.

CLARK, R. and HOUDE, J.-F. (2013). Collusion with asymmetric retailers: Evidence from a gasoline price-fixing case. *American Economic Journal: Microeconomics*, **5** (3), 97–123.

COMPETITION COMMISSION OF SINGAPORE (2009). Ccs fines 16 coach operators and association uss 1.69 million for price-fixing. Press Release.

CORTS, K. S. (1999). Conduct parameters and the measurement of market power. *Journal of Econometrics*, **88** (2), 227–250.

CRAMTON, P. and TRACY, J. (2003). Unions, bargaining and strikes. In *International Handbook of Trade and Unions*, Edward Elgar.

DONOVAN, W. J. (1926). The legality of trade associations. *Proceedings of the Academy of Political Science in the City of New York*, **11** (4), 19–26.

EIZENBERG, A. and SALVO, A. (2015). The rise of fringe competitors in the wake of an emerging middle class: An empirical analysis. *American Economic Journal: Applied Economics*, **7** (3), 85–122.

FABRA, N. (2006). Collusion with capacity constraints over the business cycle. *International Journal of Industrial Organization*, **24** (1), 69–81.

FNE (2013). Requerimiento. *Rol C No. 265-13*.

— (2014). Observaciones a la prueba. *Rol C No. 265-13*.

FTC (2018). Spotlight on trade associations. `https://www.ftc.gov/tips-advice/competition-guidance/guide-antitrust-laws/dealings-competitors/spotlight-trade`, accessed: 2018-06-30.

GANDHI, A., LU, Z. and SHI, X. (2017). Estimating demand for differentiated products with zeroes in market share data. *Working Paper*.

GENESOVE, D. and MULLIN, W. P. (1998). Testing static oligopoly models: conduct and cost in the sugar industry, 1890-1914. *The RAND Journal of Economics*, pp. 355–377.

— and — (2001). Rules, communication, and collusion: Narrative evidence from the sugar institute case. *American Economic Review*, **91** (3), 379–398.

GOWRISANKARAN, G., NEVO, A. and TOWN, R. (2015). Mergers when prices are negotiated: Evidence from the hospital industry. *American Economic Review*, **105** (1), 172–203.

GU, W. and KUHN, P. (1998). A theory of holdouts in wage bargaining. *American Economic Review*, pp. 428–449.

HARRINGTON, J. E. (2016). Heterogeneous firms can always collude on a minimum price. *Economics Letters*, **138**, 46–49.

— and SKRZYPACZ, A. (2011). Private monitoring and communication in cartels: Explaining recent collusive practices. *American Economic Review*, **101** (6), 2425–49.

HICKS, J. (1932). *The Theory of Wages*. New York: Macmillan.

HO, K. and LEE, R. S. (2017). Insurer competition in health care markets. *Econometrica*, **85** (2), 379–417.

IBANEZ, C. (2017). Caracterización del mercado de seguros complementarios de salud en base a encuesta casen 2015. Documento de Trabajo, Superintendencia de Salud.

IGAMI, M. and SUGAYA, T. (2017). Measuring the incentive to collude: The vitamin cartels, 1990-1999, working Paper.

KENNAN, J. (1986). The economics of strikes. *Handbook of labor economics*, **2**, 1091–1137.

KIRBY, A. J. (1988). Trade associations as information exchange mechanisms. *The RAND Journal of Economics*, pp. 138–146.

KÜHN, K.-U., MATUTES, C. and MOLDOVANU, B. (2001). Fighting collusion by regulating communication between firms. *Economic Policy*, **16** (32), 169–204.

LEE, R. S. and FONG, K. (2013). Markov perfect network formation: An applied framework for bilateral oligopoly and bargaining in buyer-seller networks. Working Paper.

LEVENSTEIN, M. C. (1997). Price wars and the stability of collusion: A study of the pre-World War I bromine industry. *The Journal of Industrial Economics*, **45** (2), 117–137.

— and SUSLOW, V. Y. (2006). What determines cartel success? *Journal of economic literature*, **44** (1), 43–95.

— and — (2011). Breaking up is hard to do: Determinants of cartel duration. *Journal of Law & Economics*, **54**, 455–462.

MCGAHAN, A. M. (1995). Cooperation in prices and capacities: Trade associations in brewing after repeal. *The Journal of Law and Economics*, **38** (2), 521–559.

MILLER, N. H. and WEINBERG, M. C. (2017). Understanding the price effects of the MillerCoors joint venture. *Econometrica*, **85** (6), 1763–1791.

MOTTA, M. (2004). *Competition Policy: Theory and Practice*. Cambridge University Press.

NEVO, A. (1998). Identification of the oligopoly solution concept in a differentiated-products industry. *Economics Letters*, **59** (3), 391–395.

— (2001). Measuring market power in the ready-to-eat cereal industry. *Econometrica*, **69** (2), 307–342.

OLIPHANT, H. (1926). Trade associations and the law. *Columbia Law Review*, **26**, 381.

PHELPS, C. E. and NEWHOUSE, J. P. (1974). Coinsurance, the price of time, and the demand for medical services. *the Review of Economics and Statistics*, pp. 334–342.

PORTER, R. H. (1983). A study of cartel stability: the Joint Executive Committee, 1880-1886. *The Bell Journal of Economics*, pp. 301–314.

RÖLLER, L.-H. and STEEN, F. (2006). On the workings of a cartel: Evidence from the norwegian cement industry. *American Economic Review*, **96** (1), 321–338.

STAIGER, R. W. and WOLAK, F. A. (1992). Collusive pricing with capacity constraints in the presence of demand uncertainty. *The RAND Journal of Economics*, pp. 203–220.

SULLIVAN, C. (2017). The ice cream split: Empirically distinguishing price and product space collusion, working Paper.

SYMEONIDIS, G. (2002). *The effects of competition: Cartel policy and the evolution of strategy and structure in British industry*. MIT press.

VIVES, X. (1990). Trade association disclosure rules, incentives to share information, and welfare. *the RAND Journal of Economics*, pp. 409–430.

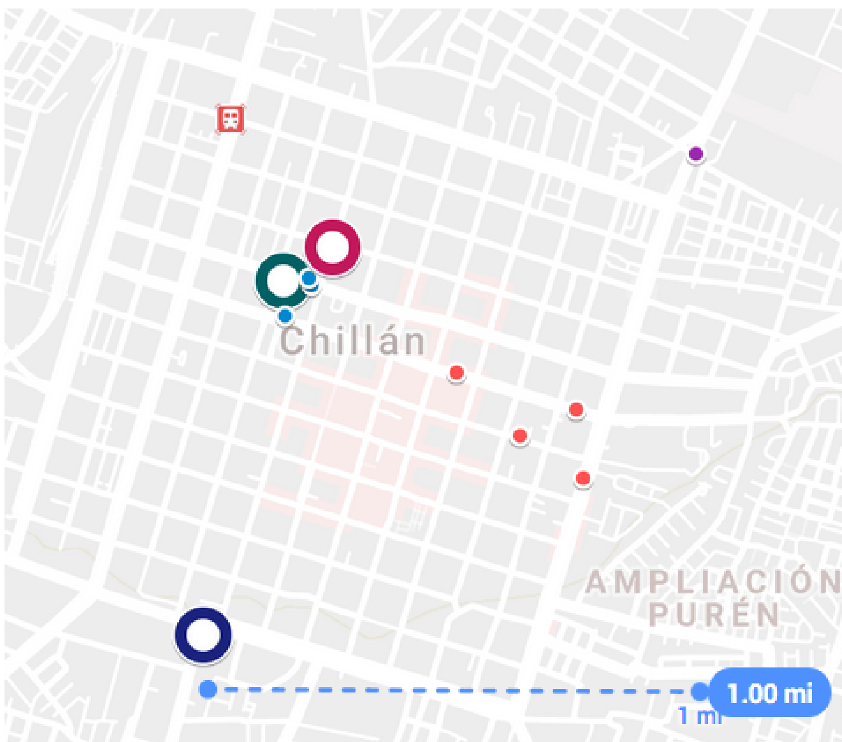
WHINSTON, M. D. (2008). *Lectures on Antitrust Economics*. The Cairoli Lectures.

# Appendix

## A1 Location and Nests

The following figure plots the geographic distribution of physicians' main address for the estimation of the nested logit demand model. The physicians' address come from FNE (2013) and, for the physicians that were not indicted, from local websites (buscachillan.cl, doctoralia.cl) or Fonasa.

**Figure A1:** Locations and Nests




Note: The Figure plots the geographic distribution of physicians' main address for the estimation of the nested logit demand model. Each color represents a different nest. Large circles represent groups of five or more co-located physicians. Each of those groups constitute a different nest. Small dots represents groups of three or less physicians, which are nested in three different nests based on the geographic distance.

## A2 Zero-share correction

The following table shows the demand estimates that result from different strategies to handle the presence of market shares equal to zero in the nested logit model. Column (1) shows the results from omitting zero shares (i.e., dropping observations with no sales). Column (2) replaces zeros with a value of 1. Finally, Column (3) shows the results from our preferred specification, which corrects the zeros following Gandhi *et al.* (2017). We find that omitting the zeros is inconsistent with utility maximization ($\sigma > 1$). We also find that replacing the zeros with ones underestimate the price elasticity, but to a lesser extent.

**Table A1:** Nested Logit Estimates–Zero Market Shares Correction

| | (1) | (2) | (3) |
|---|---|---|---|
| $p_{ijt}(1-c_{ijt})$ | -0.015 | -0.028*** | -0.050*** |
| | (0.011) | (0.008) | (0.018) |
| $\ln within-share$ | 1.066*** | 0.850*** | 0.735*** |
| | (0.148) | (0.108) | (0.152) |
| Correction | No Zeros | Ones | Dirichlet |
| $\eta_{pre}$ | 0.62 | -0.61 | -0.65 |
| $\eta_{post}$ | 1.63 | -1.58 | -1.68 |
| Observations | 4870 | 6620 | 6620 |
| AR $F$-stat | 52.04 | 73.72 | 39.42 |
| AP $F$-stat 1 | 329.78 | 429.49 | 334.84 |
| AP $F$-stat 2 | 31.90 | 46.55 | 19.83 |

Note: The Table shows the demand estimates of alternative forms for correcting for the zero market shares. All specifications include month-of-the-year and year fixed effects. Out of pocket expenditures are in thousand CLP. The AR $F$-stat corresponds to the Anderson-Rubin Wald $F$-statistic. The AP $F$-stat 1 and 2 correspond to the first-stage $F$ statistics of the excluded instruments for out-of-pocket expenditure and within shares respectively (Angrist and Pischke, 2008). Heteroscedasticity robust standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$

## A3 Reduced-form evidence of patients' response

In this section we provide descriptive evidence of significant demand responses to the large price increases documented in Section 2.2. In particular, we show that the price increases among colluded doctors lead to significant shift of patients across doctors, especially from colluding doctors towards non-colluding doctors.

We define a visit of patient $i$ to doctor $j$ to represent a "switch" whenever the doctor seen by $i$ in her previous visit to an OBGyN was $j' \neq j$. Figure A2 shows switching rates (i.e. the share of visits defined as a "switch") over time for Ñuble (the province where doctors colluded) and for a set of nearby provinces that we use as "control" provinces.[32] The figure shows that, before the agreement, the switching rates in Ñuble and control provinces have a parallel trend, and are stable over time. However, switching rates increased in Ñuble after February 2012 from 22 to 32 percent while there is no discernible increase in the switching rates in the control provinces.

We formalize the previous discussion by estimating a differences-in-differences model for switching rates. Let $w_{plt}$ equal to 1 if the visit by patient $p$ in location $l$ in period $t$ corresponds to a switch, and 0 otherwise. We estimate, by OLS, the parameters of the following equation:

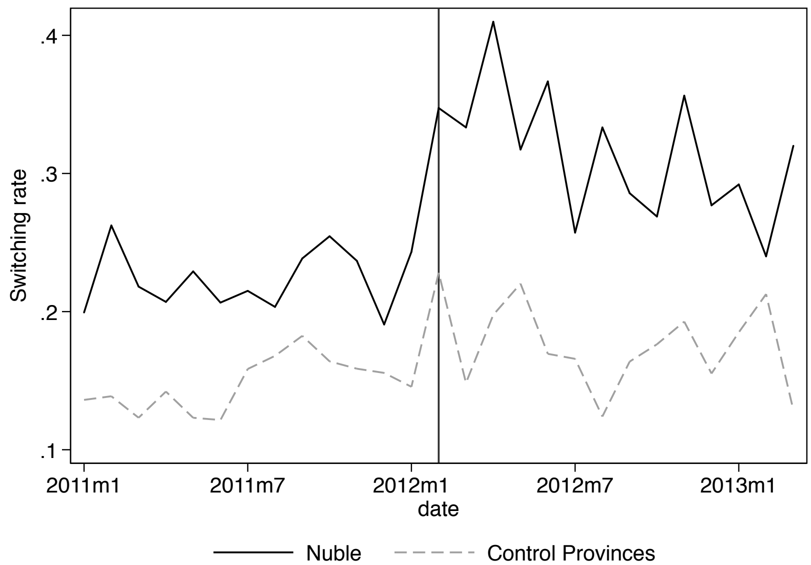
$$w_{plt} = \alpha_t + \beta T_l + \delta After_t + \gamma T_l \times After_t + \lambda_t + \epsilon_{ilt}, \tag{A1}$$

where $T_l$ is an indicator for Ñuble and $\lambda_t$ is a set of controls for calendar time. We are interested in $\gamma$, the estimated effect of the collusion on switching rates in Ñuble.

Table A2 shows the estimation results of Equation (A1). Column (1) does not

[32]We use "Concepcion" and "Biobío" as control provinces, although the results are robust to using a different subset of provinces in the region as a control group. See footnote 9 for details.

**Figure A2:** Provider switching rates in Ñuble and surrounding provinces




Note: The figure shows the average switching rates across doctors for OBGyN visits in the province of Ñuble (where collusion occurred) and in surrounding provinces (Biobío and Concepción). The vertical line marks the date of the collusion.

include time fixed effects, whereas column (2) replaces the $\delta After_t$ term by a full set of year-month fixed effects. In both specifications, we find $\hat{\gamma} = 0.07$ —indicating an increase in switching rates of 7 percentage points— corresponding to a 32 percent increase from the baseline switching rate in Ñuble before the collusion.

This large increase in switching across physicians occurred mostly from colluding towards non-colluding physicians who did not raise their prices after February 2012. Figure A3 plots the percentage changes in out-of-pocket prices against percentage changes in visits in the pre and post period (weighted by total number of visits in the period). We find that the two non-colluding doctors increased their visits by 206 percent on average, while the median colluding doctor decreased their visits by 12 percent.

**Table A2:** Doctor switching rates

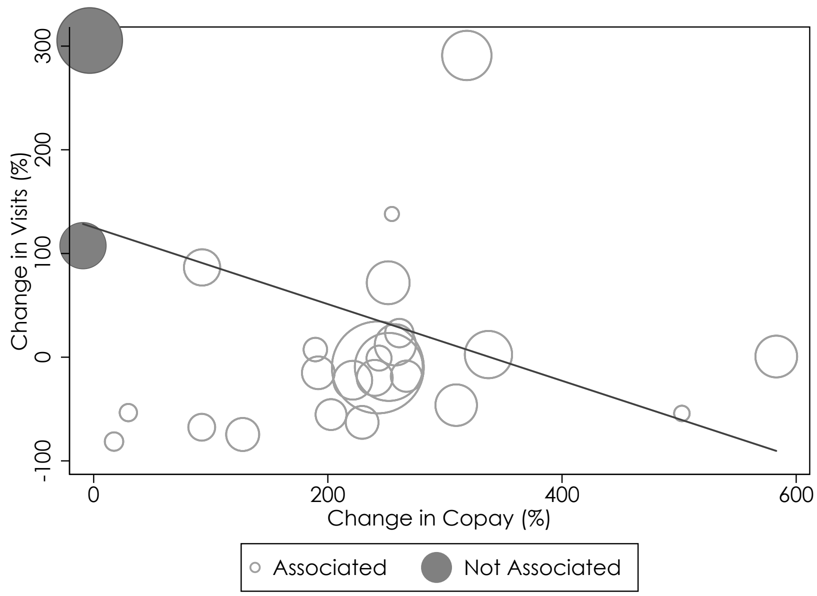
| | (1) | (2) |
|---|---|---|
| $T_l = 1$ | 0.08*** | 0.08*** |
| | (0.01) | (0.01) |
| $After_t$ | 0.03*** | |
| | (0.01) | |
| $(After_t) \times T_l$ | 0.07*** | 0.07*** |
| | (0.02) | (0.02) |
| Year-month FE | N | Y |
| $N$ | 16,742 | 16,742 |
| $R^2$ | 0.02 | 0.02 |
| Baseline switching | 0.22 | 0.22 |

Note: The baseline switching rate corresponds to the switching rate in Ñuble before the agreement period Heteroscedasticity robust standard errors in parentheses. * p<0.10, ** p<0.05, *** p<0.01

## A4 Supplemental Insurance

A share of the population in Chile buys supplemental insurance through private insurance companies to cover out-of-pocket expenditures not covered by their plan in Isapre (or FONASA). Unfortunately we do not have data on the extent of supplemental coverage held by Isapre affiliates in our dataset. If patients of colluding physicians decided to buy supplemental insurance to cover the higher out-of-pocket expenses in gynecologist visits, our demand elasticity would be biased downwards. To alleviate this concern, we use Chile's National Socioeconomic Characterization Survey (CASEN) to show that the rate of individuals who respond to hold supplemental coverage did not change diferentially after the collusion in the region where doctors colluded. The CASEN survey asks households heads about their primary source of insurance (Isapre, FONASA, or other) and whether

**Figure A3:** Changes in Market Share and Changes in Out-of-Pocket Prices




Note: The Figure plots the percentage changes in out-of-pocket prices against percentage changes in visits in 2012 with respect to 2011 for every physician, and a linear fit. The size of the markers correspond to the physician's total number of visits in the period, which are also the linear fit weights.

any member in their household holds supplemental insurance.

Using the 2011 and 2013 waves of the survey we run a differences-in-differences regression similar to (A1) for supplemental coverage on the sub-sample of Isapre affiliates in Nuble and control provinces. The results from this exercise are shown in Table A3.

The differences-in-differences coefficient is negative but not statistically significant, which implies that there is no evidence that supplemental coverage differentially increased in Nuble between 2011 and 2013. [33]

[33] We note that the share of privately insured individuals holding supplemental coverage as declared in the CASEN survey is large (35 percent in 2011). One potential explanation is the fact that individuals may respond "yes" if they purchased any add-on to their Isapre plan instead of supplemental coverage as noted in Ibanez (2017). In fact, using the 2009 wave of the Social Protection Survey we find that only 10 percent of respondents declare to have supplemental coverage. The main drawback of using the Social Protection Survey for this analysis is that it does not contain information of the district of residence, and is not available in years close to our period of study

**Table A3:** Supplemental Insurance Coverage

| | (1) |
|---|---|
| | Supplemental Insurance |
| $t = 2013$ | 0.05 |
| | (0.03) |
| $T_l = 1$ | -0.02 |
| | (0.05) |
| $T_l \times (t = 2013)$ | -0.06 |
| | (0.07) |
| constant | 0.35*** |
| | (0.02) |
| $N$ | 1,196 |
| $R^2$ | 0.00 |

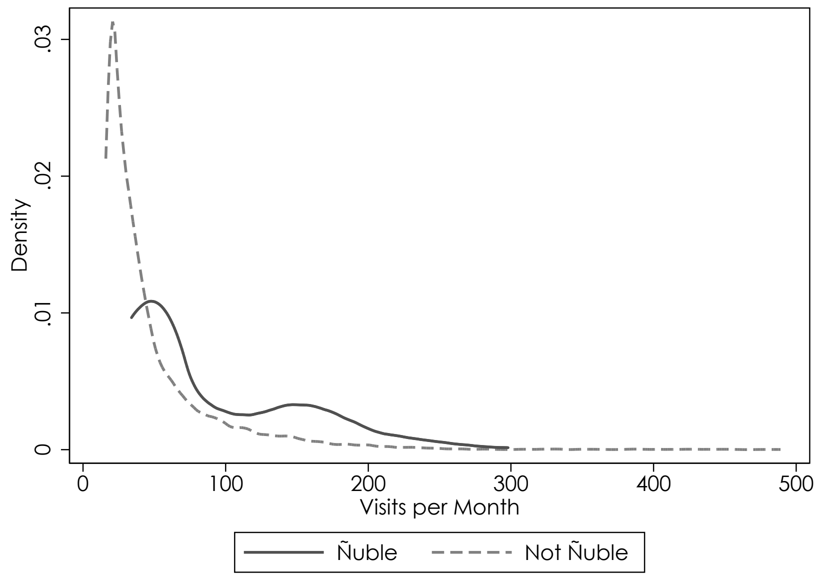
Standard errors in parentheses
* p<0.10, ** p<0.05, *** p<0.01

## A5 Capacity Constraints

In this Section we provide descriptive evidence that indicates no observable (binding) capacity constraints. First, we plot the right tail of the distribution of visits per month of ob/gyns in Ñuble and in other neighboring provinces in Figure A4. Specifically, we plot the distribution of visits between the 75 and the 99.9 percentile of the visits distribution (we right-censor the distribution due to possible measurement error). We find that visits to ob/gyns outside Ñuble have a longer right tail. Assuming that capacity constraints are similar across locations, the figure suggests that capacity contraints are not binding in Ñuble.

Second, we look at the predicted quantities in the Nash supply model. In the pre-agreement period, the maximum predicted quantity is 124 visits, and in the post-agreement period the maximum predicted quantity is 140. To make sense

**Figure A4:** Right Tail of the Visits per Month Distribution




Note: The Figure plots the distribution of the 75-99.9 percentiles of the visits per month distribution of ob/gyn in Ñuble and in other neighboring provinces.

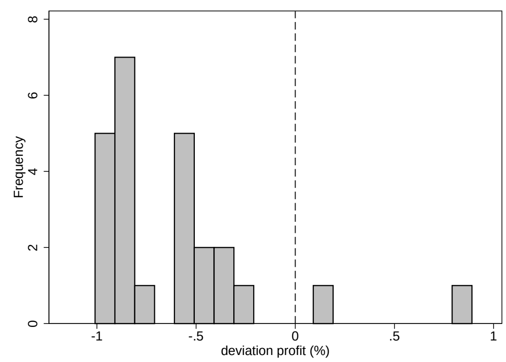
of these numbers, the average ob/gyn visit in the US lasts roughly 20 minutes.[34] Hence, a doctor working an 8-hour shift, 5 days a week, is able to see up to 480 patients. Therefore, the predicted number of visits are significantly smaller than a doctor's maximum capacity. The low predicted (and actual) number visits is due to the fact that we are only considering visits from the private sector. If private visits are more profitable than public patients, doctors may switch towards seeing private patients before their capacity constraint becomes binding.

[34]See "Profile of Ob-Gyn Practice 1991-2003", *American College of Obstetricians and Gynecologists*, `https://www.acog.org/~/media/Departments/Practice/ProfileofOb-gynPractice1991-2003.pdf`. Accessed Oct 25, 2018.

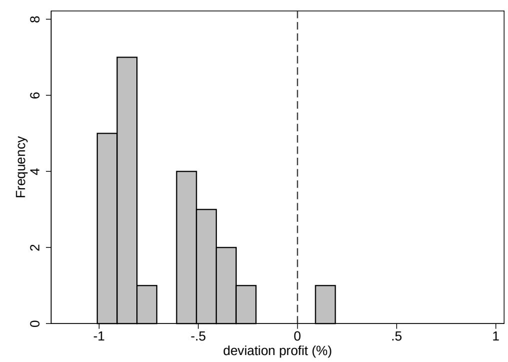
# A6 Incentives to Join

Figure A5 shows the resulting deviation profits after considering endogenous repricing, discussed in the last paragraph of section 5.3. With endogenous repricing, we find that only two out of the 25 physicians had incentives to remain out-of-network and not join the Association. This result mirrors the results of Figure 6 in the main text (where no endogenous repricing is allowed) because the change in the Associated physicians' prices when one single physician leaves the Association is small.

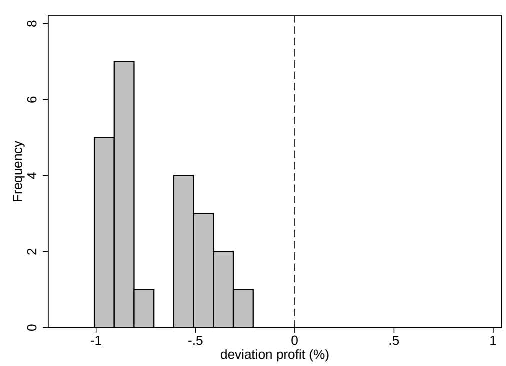
**Figure A5:** Deviation Profits and Association Stability




**(a)** Step 1




**(b)** Step 2




**(c)** Step 3

The Figure shows the distribution of per-period deviation profits across physicians. Each point in the figure corresponds to the deviation profit of a different physician, as a percentage of the profit from staying in the Association.