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

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Abstract

We study a recent case where 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 shed light on the role of trade association in collusion among a large number of heterogeneous agents, and provide insights for the antitrust analysis of trade associations.

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

JEL classification: I11, L13, L41

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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, OECD, 2007, 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 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. The setting—a trade association of gynecologists in Chile—mirrors several cases challenged by the FTC in which a large physician group colluded in negotiations with payers.[1]

The Association started working in mid-2011 and comprised 90 percent of the local gynecologists in one city. After six months of unsuccessful negotiations with the insurance companies to increase prices, the members of the Association simultaneously terminated their contracts with the insurers and agreed to set their fees above a specified minimum price. As a result of those two measures, all associated physicians became out-of-network providers and out-of-pocket visit prices rose almost 200 percent on average. The Association operated for 28 months before it was challenged by the National Economic Prosecutor and was ultimately abolished by the Supreme Court for collusive practices

The rich data available for this case provides a unique opportunity to better understand the emergence and antitrust consequences of trade associations. The goal of our empirical analysis is to provide a comprehensive framework to analyze the gynecologists case through the lens of antitrust. We rationalize the Association's pricing strategy, assess its stability, and evaluate the role of coordination in the failure of the negotiations with insurers.

[1] See `https://www.modernhealthcare.com/article/20090112/MODERNPHYSICIAN/301049974/ftc-keeps-wary-eye-on-network-arrangements` and Meier *et al.* (2017). Also, the prosecution of this case was heavily based on the treatment of similar cases by the FTC.

We start by showing reduced-form evidence that the Association brought about large and unexpected increases in out-of-pocket prices, and that these resulted in increased rates of patient switching across doctors. We then use a demand model to estimate patients' price elasticities for visits. The implied price elasticities serve as an input of a supply-side model that we calibrate to recover the degree of price coordination that better fits the prices set by the Association.

Our structural analysis shows that the realized prices of the physicians in the Association coincided with Nash-Bertrand prices. Moreover, Nash prices were such 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.

In light of our results, we use our estimated model to empirically investigate the incentive compatibility of the Association's strategy. We calculate 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 Association's 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 demand and supply models allow us to investigate the role of the Association in the negotiation process. To this aim, we compare the profits from the decision to leave the network jointly to the counterfactual profits from leaving the network unilaterally. We find that most doctors would have benefited from unilaterally terminating their contracts with the insurers. Yet, the fact that doctors did not leave the network unilaterally is consistent with the case documents showing that the physicians' goal was to reach ultimately better terms with the insurers. In addition, for every doctor, the profits under the Association are substantially larger than those under unilateral contract termination: On average, leaving the insurers jointly results in profits that are 2.5 times larger than leaving unilaterally. This large gap helps rationalizing why the Association triggered the breakdown of the bargaining between doctors and insurance companies. The Association created an outside option with profits large enough that there was no longer surplus created by the agreement with insurers.

Our results have implications for the antitrust analysis of trade associations in the context of vertical negotiations. As mentioned, antitrust offences by physicians associations are quite common. In fact, the FTC has brought more than 35 complaints against associations of independent physicians since 2001, most of which involved a large group of physicians that terminated their agreement with insurers in the context of a negotiation with payers.[2] However, the position taken by the FTC in these cases is not without controversy: Recent calls by antitrust practitioners put into question the prosecution of agreements that seek to counteract large imbalances of power in negotiations, including not only physicians in their negotiations with insurers, but also hotels in their negotiations with online platforms and truck owners in their negotiations with transport companies (see, e.g., Mullan and Timan, 2018; ACCC, 2018). Moreover, Alaska, New Jersey, and Texas have passed bills that permit physicians to bargain collectively (Blair and Herndon, 2004). More recently, the DOJ and FTC have explicitly allowed joint negotiations with payers among providers that get together as "Accountable Care Organizations" (ACO) to participate in the Medicare Shared Savings Program, deeming joint negotiations as "reasonably necessary to an ACO's primary purpose of improving health care delivery" (DOJ and FTC, 2011).[3].

Our estimates show that the bargaining with insurance companies significantly restrained physician prices in the pre-association period. Forming the Association helped its members leave a bargaining equilibrium that entailed low profits for them, but did not result in collusive prices. Still, the Association decreased consumer welfare by appropriating part of the surplus that was captured by insurers and their clients through lower visit prices.

Our analysis also provides several insights for our understanding of collusion. A first insight is related to the role of communication. Standard economic models and conventional wisdom predict that price coordination in a large and heterogeneous set of firms is hard (e.g., Motta, 2004; Levenstein and Suslow, 2006). We find evidence for this prediction in that the trade association was not effective in raising prices above competitive levels despite sustained communication. However, coordination was effective to terminate the contracts with insurers, which is arguably some-

[2]In these cases the FTC generally issued a consent order that prohibits Associations to negotiate on behalf of any physician with any payer, to refuse to deal with any payer, or to have involvement into the individual negotiations between doctors and payers

[3]The high degree of concentration makes the health insurance industry a natural market for the emergence of proposals seeking to counteract insurer bargaining power. Fulton (2017) finds that 57 percent of health insurance markets were highly concentrated in 2016. The American Medical Association reports that 69 percent were highly concentrated ((Association, 2011))

thing less complex than coordinating on prices. These facts constitute a challenge for theorists in understanding the different roles of communication in cases of collusion.

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 according to antitrust law (see Whinston, 2008 for a discussion).

Finally, we show evidence from a case that is consistent with Harrington's (2016) insight that there always exists a minimum price at which firms can successfully collude in differentiated-product industries. Our analysis permits to evaluate the level of the minimum price chosen in reality against the range of levels theoretically predicted by Harrington (2016). The use of a minimum price in the gynecologists case is consistent with Harrington's claim that a minimum price is indeed a device that heterogeneous firms may use to collude. However, in this case the minimum price was ineffective because it was not binding at the competitive price level.

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).[4] 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. We also study empirically the role of a minimum price.[5]

Few papers examine collusion on dimensions other than price. An exception is Sullivan (2017) that reports collusion on product characteristics. We analyze collusion on the organization of the vertical network relationship. The vertical structure determined consumers' out-of-pocket

[4]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.

[5]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).

expenses, and although the network is related to price setting, it is also different from it. Thus, in our counterfactuals doctors choose a share of copay and a price, which constitute two different decision variables with respect to which doctors maximize profits. We find that coordination on contract termination was more important than price coordination in increasing profits.

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 the prices set by the Association after the breakdown of the negotiations with the insurers are best predicted by Nash-Bertrand prices rather than any positive degree of price coordination.

This paper is also related to the literature on the anticompetitive effects of trade associations.[6] 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 associations before antitrust legislation.[7] 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 (Lee and Fong, 2013).

Some papers in the antitrust literature on health care document higher prices in areas where physician practices face less competition (Dunn and Shapiro 2014; Baker *et al.* 2014; Austin and

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

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

Baker 2015). Our paper shows explicitly that physician coordination may lead to higher prices and that bargaining is an important mechanism determining consumer prices for doctor visits.

We also contribute to the literature on the antitrust treatment of bargaining in the context of vertical negotiations. Blair and Herndon (2004), Blair and DePasquale (2011), Campbell (2007, 2008) and Baker *et al.* (2008) theoretically discuss whether upstream providers negotiating jointly with downstream firms may improve welfare. Interestingly, this theoretical literature is highly motivated by the physician-insurer cases.[8]

A strand of the literature in labor economics studies strikes and labor disputes. Classic references on collective bargaining and strikes are Hicks (1932) and Ashenfelter and Johnson (1969).[9] The main feature that differentiates a trade association from a labor union is that employees are not considered economic entities because they do not bear the risk of the commercial activity. Hence, a labor union that purely represents employees is not considered an association for the purposes of antitrust law.[10] Moreover, workers do not play a "wage game" similar to the pricing game played by physicians of the association.[11]

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 by monthly contributions deducted from labor income, cost-sharing, and resources from the general government. FONASA covers roughly two thirds of the population

[8] Blair and Herndon (2004) refers to the physician-insurer case explicitly, while Campbell (2008) and Blair and DePasquale (2011) also motivate their analysis with the physician-insurer case.

[9] Card (1990), Gu and Kuhn (1998), and Cramton and Tracy (2003) provide theoretical models and empirical evidence. Kennan (1986) and Cramton and Tracy (2003) review this literature.

[10] OECD (2007), p. 21.

[11] We thank an anonymous referee for making this point.

(about 11 million people).[12]

Plans in the private system are comprehensive, stand-alone plans. These plan have two main coverage features: coinsurance rates (one for inpatient care and another for outpatient care) and coverage caps (insurer payment caps).[13] Similarly to private plans in the U.S., 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 lower reimbursement rate (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 US. 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. As is common in health care markets, in-network private providers negotiate their rates with insurance companies through bilateral negotiations On the other hand, out-of-network private providers set their rates privately.[14] Finally, the public payer (FONASA) reimburses private providers following a pre-determined rate for each procedure and a provider quality rank. [15]

2.2 The Antitrust Case

Gynecologists in Chillán, a city of 175,000 inhabitants in the Ñuble province in southern Chile, legally constituted the Gynecologists' Association of Ñuble in August 2011.[16] The Association

[12]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).

[13]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.

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

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

[16] 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".

was formed by 26 out of the 29 ob/gyns of Chillán.[17]

Upon the creation of the Association, all gynecologists in Ñuble held network contracts with Isapres that were negotiated independently. To improve their bargaining position vis-à-vis the insurance companies, members of the association started to jointly negotiate their fees with Isapres in 2011.[18] 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).[19] 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.[20] 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 complaints, but did not accept the physicians' requests immediately.[21]

Throughout 2012 the insurance companies contacted either the physicians separately or the Associaton's head and offered rates lower than those requested by the Association. The physicians rejected those private offers (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). [22] It was not until March 2013 that a major Isapre accepted (almost fully) the Association's terms.[23] Yet, that same month the FNE started

[17]One of associated physicians was not indicted. However, we observe an increase in his visit prices and, thus, we consider him part of the collusive group.

[18]Bargaining is frequently given as the main motive for trade associations in similar antitrust cases. For example, a physician association in South Africa claimed that physicians "needed [the association] to balance the strong bargaining power of private hospitals." (OECD, 2007)

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

[20]This course of action was recorded in an official act, which also stated that the measures were adopted because the members "feel that they are being exploited" by the Isapres, which "have not known how to compensate and value" the physicians' work. (FNE, 2014, p. 53).

[21]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.

[22]For example, in January 2012, Dr. B. received an offer to the Association of CLP 16,000. Another physician who got a private offer answered "I feel that it would only be fair if this offer is extended to all the qualified professionals in Ñuble" (FNE, 2014 p. 68).

[23]The agreed terms were CLP 25,000 for visits and 4 times the rate paid by FONASA for surgeries. In addition, a

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 75 percent when the Association formed, from roughly CLP 13,900 in December 2011 to CLP 24,300 in February of 2012. 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 68 percent on the average coinsurance rate, from 25 percent to 42 percent. Combined, the out-of-pocket cost of a visit to the average doctor in Chillán increased by almost 200%. There were no discernible changes in the out-of-pocket costs in neighboring cities. [24,25]

[INSERT FIGURE 1 HERE]

Compliance within the Association with the minimum visit price was substantial. 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.[26]

[INSERT FIGURE 2 HERE]

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.

[24] Table A1 shows the result of estimating a differences in differences model for prices and coinsurance rates. The estimated price increase is 60% and the estimated increase in the coinsurance rate is 61% of the baseline levels in Chillán. The difference between the increase we report in the text and the estimated effect in the differences-in-differences is due to a slight upward trend in the prices in the control group.

[25] 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. Patients were surprised of the change, which was reported in the local media (FNE, 2013, p.9)

[26] 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.

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 operating in Chillán and its neighboring provinces.[27] 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 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. These two datasets combined allow us to reconstruct the universe of visits to gynecologists in Chillán and its neighboring cities.[28] Given that 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. There are 29 different physicians in the data. We drop observations corresponding to two physicians with less than 200 combined visits. Hence, there are 25 associated and 2 non-associated physicians in our final sample.[29] Finally, we obtain the physicians' main address from the case documents (FNE, 2013).[30]

Figure 3 shows the evolution of prices and profits over time for associated and non-associated physicians, as well as for the average doctor. The figure assumes that physicians' marginal cost is equal to the FONASA rate, an assumption that we discuss in more detail in Section 5.2. 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.

The profits of associated and non-associated doctors were stable before the Association, and similar across groups. However, profits changed drastically for both groups after the Association

[27]There is no entry or exit of doctors during the time period we study.

[28]We cannot match the two datasets at the patient level. For the demand estimation, we assume that the visits in the receipts database were reimbursed at the out-of-network coverage rate.

[29]In addition, we drop the physician-month observations corresponding to four physicians 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.

[30]For the physicians who were not indicted, we use local websites.

was formed. 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.

[INSERT FIGURE 3 HERE]

Figure 4 shows the evolution of total insurer costs and out-of-pocket expenses (the sum of which corresponds to physician's revenues). The figures shows that higher physician revenues came at the expense of higher costs for both patients and insurers. For insurers, this means that the lower out-of-pocket coverage and any endogenous response in demand did not compensate for the higher list price.

[INSERT FIGURE 4 HERE]

4 Demand

4.1 Descriptive Evidence on Physician Switching

We start by providing 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 a significant shift of patients across doctors, especially from colluding doctors towards non-colluding doctors.

We define a visit of patient p to doctor i to represent a "switch" whenever the doctor seen by p in her previous visit was $i' \neq i$. Panel (a) of Figure 5 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.[31] Panel (a) also shows that switching rates in Ñuble and control provinces had a parallel trend and were stable over time before the agreement. However, switching rates increased in Ñuble after February 2012 from 22 to 32 percent while there was no discernible increase in the switching rates in the control provinces. Panel (b) shows regression results of switching on a Ñuble indicator for each quarter separately. The panel

[31] 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 16 for details.

indicates that the difference between the two groups is statistically significant immediately after the association implemented the price hikes.

[INSERT FIGURE 5 HERE]

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

$$w_{pilt} = \alpha_t + \beta T_l + \gamma T_l \times After_t + g(t) + \epsilon_{ilt}, \tag{1}$$

where T_l is an indicator for Ñuble and $g(t)$ controls for calendar time. We are interested in γ, the estimated effect of the collusion on switching rates in Ñuble.

Table 1 shows the estimation results of Equation (1). Column (1) uses $g(t) = \delta After_t$, where $After_t$ is dummy variable indicating post-Association dates. Column (2) replaces $\delta After_t$ by a full set of year-month fixed effects. In both specifications we find $\hat{\gamma} = 0.07$; a 7 p.p. increase in the switching rate. This increase corresponds to a 32 percent of the baseline switching rate in Ñuble before the collusion. Appendix 5 shows that this large increase in switching across physicians occurred mostly from colluding physicians towards non-colluding physicians.

[INSERT TABLE 1 HERE]

4.2 Demand Model

We model physicians as differentiated-product firms to allow for idiosyncratic preferences for ob/gyns, which is likely an important feature of the industry. In our context, physicians are different from each other in terms of location, and possibly in terms of their quality, which introduces both horizontal and vertical differentiation.

We assume a nested logit demand model, where the set of physicians is partitioned into non-overlapping nests B_k, $k = 1, \ldots, K$.[32]

[32] The nests are determined based on the doctor's geographic location as explained in detail in Section 4.3. Using nests captures differentiation across doctors arising from patient's heterogeneous preferences for location.

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

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

where δ_{ijt} denotes the mean utility of insurees who visited doctor i at time t, ϵ_{ijt} is the unobserved mean-utility component of physician desirability, η_{pkt} is an unobserved common shock to physicians in nest k, σ is a parameter that measures within nest correlation, and ν_{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 insurance company j), follows the standard share equation:

$$s_{ijt} = \frac{\exp(\frac{\delta_{ijt}}{1-\sigma}) \sum_{h\in B_k} \left[\exp(\frac{\delta_{hjt}}{1-\sigma})\right]^{-\sigma}}{1 + \sum_{l=1}^{K} \sum_{h\in B_l} \left[\exp(\frac{\delta_{hjt}}{1-\sigma})\right]^{1-\sigma}}. \tag{2}$$

We parameterize δ_{ijt} as

$$\delta_{ijt} = \alpha p_{ijt}(1-c_{ijt}) + \mu_{ij} + f(t), \tag{3}$$

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, μ_{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.

4.3 Estimation

The estimation of Equation (2) poses three challenges. First, 26 percent of market-share observations in our data are equal to zero. This fact generates censoring in the estimation of Equation (2) because the logit model does not allow for zero shares. Second, we need to specify the nesting structure of the model. Finally, we face the standard endogeneity problem of prices and within shares, as both ϵ_{ijt} and the within shares are potentially correlated with prices. We address these issues below.

Zero Shares As mentioned, many of the doctor shares in the data are zero. However, all doctors are in all insurees' choice set because insurees can in principle see any given doctor by paying the corresponding out-of-pocket price (see Section 2). Therefore, the observed zero shares s_{ijt} can be interpreted as the realization of small shares with a finite number of consumers, which differ from the true population probabilities implied by the model.

We rely on Gandhi *et al.* (2014), who propose an asymptotic correction based on the assumption that sales distribute according to Zipf's Law.[33] In that case, then 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 (assumed to be equal to the number of female insurees in each Isapre), N represents the number of doctors, and $DCM(\cdot)$ is the Dirichlet Multinomial Distribution with the same N parameters θ_j and s_0 represents the share of the outside option.

Denote the digamma function by $\psi(\cdot)$. We invert the shares equation, as in Berry (1994), and take expectations to obtain:

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

which is a corrected version of the nested logit demand model that we can estimate. In this equation, the left hand side corresponds to the expectation of the asymptotic dependent variable of the nested logit model, $E\left[\ln(\pi_{ijt}/s_{0jt})\right]$, where π_{ijt} denotes the true shares. The term in square brackets is the expectation of the log asymptotic within share of physician i, $E\left[\ln \pi_{ijt/B_k}\right] = E\left[\ln(\pi_{ijt}/\sum_{h \in B_k} \pi_{iht})\right]$, $j \in B_k$. Our main objects of interest are the parameters α and σ.

Our estimation procedure has two steps. First, we estimate the DCM parameters θ_j of (4) *via* Maximum Likelihood. Then, we use the resulting estimates to estimate the parameters of Equation (5) using a 2-stage least squares estimator, that we discuss after we introduce the nesting structure below.

[33] Quan and Williams (2018) also implements the same correction.

Nesting Patients' distaste for travel distance is a well-documented feature of demand for medical services (see, e.g., Phelps and Newhouse, 1974). In the absence of data on patient's location, we use the nesting structure to accommodate for the fact that patients are more likely to substitute to a physician that is either in the same medical center or has a practice in a nearby location. In practice, 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.[34] Also, we construct two additional nests, one for physicians outside Chillán, and another one for the outside option.

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 price and within share variables and the unobserved demand shocks ϵ_{ijt}, as is also the case in the standard nested logit model. 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 neighboring cities before the Association emerged in Chillán. Thus, the data suggests that the emergence of the Association in Chillán was not a result of particular conditions in this market. In particular it does not seem that doctor's bargaining position *vis-à-vis* insurance companies was substantially lower in Chillán than in other neighboring cities.

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.

[34] Prices in the two types of groups are similar.For instance, the median (mean) price in the pre-Association period was 13,000 (13,226) and 13,044 (14,459) in these small and large practices, respectively.

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.[35] Hence, we identify the elasticities using changes in prices and coinsurance rates within the doctor-insurer pair before and after the price change.[36]

4.4 Demand Estimation Results

Table 2 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).[37]

In our main specification (Column (4) of Table 2) we find $\hat{\alpha} = -0.050(0.018)$ and $\hat{\sigma} = 0.735(0.152)$.[38] 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 elasticity in the period before and after the Association was formed.[39]

[INSERT TABLE 2 HERE]

[35]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.

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

[37]Appendix A3 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.* (2014) for a discussion on the drawbacks of those methods.

[38]The estimate $\sigma \in [0, 1]$, is consistent with utility maximization.

[39]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).

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.[40] This allows us to calculate competitive and collusive prices given the demand estimates.

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 $\mathbf{p^N}$ the vector with the stacked prices of all the colluding physicians, $\mathbf{c}_j$ the vector of stacked coverage rates (1 minus coinsurance rates) in insurer j and the corresponding vector of monthly visits $\mathbf{q}_j(\mathbf{p^N}, \mathbf{c}_j)$. The first-order condition of colluding physicians can be written in matrix form for each period as

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

where the operator $*$ denotes element-by-element multiplication.[41] 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. The matrix C_j represents 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.[42]

We also define the ownership matrix Ω^* for the colluding physicians. The ownership ma-

[40] We do not model bargaining with the insurers during the agreement period because physicians terminated their contracts with the insurers.

[41] We omit time indexes to simplify notation.

[42] We are making the implicit assumption that doctors can always substitute a private patient for a public patient. This assumption is justified by the well-documented lack of specialist doctors and long waiting times in the public system. As an example, in 2017 there were almost 70 thousand individuals nationwide on the waiting list to visit an ob/gyn, among which almost 50 thousand had been waiting more than 4 months (INDH, 2018).

trix specifies the degree to which each colluding physician internalizes 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 Ω^* becomes the identity matrix, which determines the Nash equilibrium; any $\kappa > 0$ determines a different partial collusive equilibrium, with a higher κ indicating a higher degree of coordination on prices. The extreme of $\kappa = 1$ corresponds to full collusion on prices.[43] For any given assumed value of κ, we solve for $\mathbf{p}^N$ by numerically finding the fixed point in Equation (6).

Capacity Constraints Physicians see a mix of patients that are either privately or publicly insured. Our supply and demand models assume that capacity constraints for visits of privately-insured patients do not bind – so that physicians can always see an extra private patient at the expense of not seeing a public patient. Binding capacity constraints in the data would imply that we underestimate the price coefficient in the demand model. Moreover, binding capacity constraints could generate corner solutions in the supply model, and reduce the potential deviation profits in the counterfactuals we study below (Staiger and Wolak, 1992).We provide evidence supporting the absence of binding capacity constraints for private visits, both in the data and in the predictions of the model, in Appendix A6.

5.2 Results

The main results of the predicted price from Equation (6) are in Figure 6. The figure presents the time series of equilibrium prices assuming Nash-Bertrand competition among associated doctors

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

($\kappa = 0$).[44] 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 7 shows the distribution of average Nash prices of each physician during the Association period.

[INSERT FIGURE 6 HERE]

[INSERT FIGURE 7 HERE]

Three facts about our results are particularly noteworthy. First, as shown in Figure 6, 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.

Figure 6 also shows that the average Nash-Bertrand price is almost equal to average price set by the members of the Association. By contrast, the collusive price (when $\kappa = 1$) is equal to CLP 52,336; twice as high as the average observed price. Panel (a) of Figure 8 shows the ratio between the (quantity-weighted) average predicted price and (quantity-weighted) average observed price. Also, $\kappa = 0$ minimizes the difference between the prices predicted by the supply model and the true prices. Panel (b) of Figure 8 shows the (quantity-weighted) average of absolute differences between the predicted and observed prices is increasing in κ, formally defined as

$$f(\kappa) \equiv \quad \frac{1}{T} \sum_{t \geq t_0} \frac{\frac{1}{N} \sum_i |1 - \frac{P^N_{it}(\kappa)}{P_{it}}| q_{ijt}(\kappa)}{\sum_i q_{ijt}(\kappa)},$$

where t_0 is the date when the Association was formed, T is the number of months post-association. P_{it} is the observed price and $P^N_{it}(\kappa)$ and $q_{ijt}(\kappa)$ are the predicted prices and quantities as a function of κ. As shown in Panel (b) of Figure 8, $f(\kappa)$ is increasing in κ so that the no-coordination solution provides the best fit to the data.[45]

[44] Because 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.

[45] $f(0) = 0.17\%$, i.e., our prediction error is approximately 17 percent of the actual price.

Third, as shown in Figure 7, the distribution of Nash-Bertrand prices is such that the minimum price barely binds.

[INSERT FIGURE 8 HERE]

Our empirical 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 from the fact that the physicians were no longer constrained by their previous agreements with the insurers. Second, our findings allow us to assess the role of the minimum price as an effective collusive mechanism. In particular, Harrington (2016) shows that the use of a minimum price is an incentive compatible way of achieving *supra-competitive* prices for heterogeneous firms. Yet, our results show that the level chosen for the minimum price was not binding in the Nash equilibrium. Therefore, the minimum price was not helpful to raise profits above competitive levels. Hence, we interpret the use of a minimum price as a focal price, which helped doctors coordinate into switching to the competitive equilibrium.

Robustness to the marginal cost assumption An important assumption in our empirical analysis is the level of the marginal cost, which we assumed to be equal to the FONASA rate. However, 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 FONASA rate (CLP $10,970$) would correspond to a lower bound for the marginal cost. On the other hand, the actual pre-agreement price (CLP $13,982$) is an upper bound on the marginal cost, as a higher marginal cost would violate doctors' participation constraint in the agreement with insurers. In the extreme case that the pre-agreement price was equal to the doctor's marginal cost —a situation in which the pre-association price leaves doctors just indifferent between participating in the market or not— we would underestimate the equilibrium prices by CLP $3,000$. Still, the conclusion about lack of price coordination is unaffected by assuming a higher marginal cost.

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 price at their pre-Association average, that is, we assume they return to their original insurer contract.

Formally, for a set $\mathcal{S}$ of doctors in the Association, the vector of prices $\mathbf{p}(\mathcal{S})$ has element i defined by:

$$p_i(\mathcal{S}) = \begin{cases} p_i^d, & \text{if } i \notin \mathcal{S} \\ p_i^N, & \text{if } i \in \mathcal{S} \end{cases}$$

where p_i^N is the ith element of the vector defined implicitly by Equation 6 for doctors in $\mathcal{S}$, and p_i^d is the average of in-network prices of doctor i.

Similarly, the vector of coinsurance rates ($\mathbf{c}_j$) has element i defined by:

$$c_{i,j}(\mathcal{S}) = \begin{cases} c_{i,j}^I, & \text{if } i \notin \mathcal{S} \\ c_{i,j}^O, & \text{if } i \in \mathcal{S} \end{cases}$$

where $c_{i,j}^I$ an $c_{i,j}^O$ denote the in-network and out-of-network coinsurance rates, respectively

The profits for physician i under $\mathcal{S}$ depends on physician i's price and coinsurance rate as well as all other physician's prices and coinsurance rates, denoted by $\mathbf{p}_{-i}(S)$ and $\mathbf{c}_{-i}(S)$;

$$\pi_i(\mathcal{S}) = \sum_j (p_i(S) - mc)\, q_j \left(p_i(S), \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}(\mathcal{S}), \mathbf{c}_{-i,j}(\mathcal{S})\right).$$

When physician i deviates from the Association her deviation profits are

$$\begin{aligned} \pi_i(\mathcal{S} \setminus i) - \pi(\mathcal{S}) = & \sum_j \left(p_i^d - mc\right) q_j \left(p_i^N, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^I, \mathbf{c}_{-i,j}(\mathcal{S})\right) \\ & - \sum_j \left(p_i^N - mc\right) q_j \left(p_i^d, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^O, \mathbf{c}_{-i,j}(\mathcal{S})\right) \end{aligned} \tag{8}$$

When 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. These two effects can be decomposed by re-writing equation (8) as

$$\begin{aligned} & \pi_i(\mathcal{S} \setminus i) - \pi_i(\mathcal{S}) = \\ & \sum_j \left(p_i^d - mc\right) q_j \left(p_i^d, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^O, \mathbf{c}_{-i,j}(\mathcal{S})\right) - \sum_j \left(p_i^N - mc\right) q_j \left(p_i^N, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^O, \mathbf{c}_{-i,j}(\mathcal{S})\right) \\ & + \sum_j \left(p_i^d - mc\right) \left(q_j \left(p_i^d, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^I, \mathbf{c}_{-i,j}(\mathcal{S})\right) - q_j \left(p_i^d, \mathbf{p}_{-i}(\mathcal{S}), c_{i,j}^0, \mathbf{c}_{-i,j}(\mathcal{S})\right)\right) \end{aligned} \tag{9}$$

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. As 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.[46] Therefore, deviation profits have an ambiguous sign, so we quantify them empirically using our demand estimates.

[INSERT FIGURE 9 HERE]

Panel (a) of Figure 9 presents the results. The panel shows the histogram of deviation profits

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

across physicians, calculated as a share of the collusive profits, $\left(\pi_i(\mathcal{S}) - \pi_i(\mathcal{S} \setminus i)\right) / \pi_i(\mathcal{S})$. Each unit of observation corresponds to a different deviating physician (and, therefore, to a different counterfactual scenario) among the 25 colluding physicians in our sample.

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 leave the Association as it was actually formed. Panel (b) o Figure 9 shows the distribution of 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 under the assumption that everyone else joins and charges prices $\mathbf{p}_{-i}\left(\mathcal{S}\right)$. 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, $\tilde{\pi}\left(\mathcal{S} \setminus i\right)$, that depend on the outcome of the Nash-Bertrand pricing game when physician i does not join, $\mathbf{p}_{-i}\left(\mathcal{S} \setminus i\right)$.

We show the resulting deviation profits after considering endogenous repricing of the associated physicians in Appendix A7 . We find that price adjustments are small so that deviation profits in this case are similar to the results of Figure 9. We interpret this finding as showing that joining the Association was incentive compatible.

5.4 The role of coordination

In this section we calculate every doctor's profits in the counterfactual scenario where she would have unilaterally left the network. In this exercise, the focal doctor i sets her visit price according to the first order condition in Equation (6), while the others keep their prices at their individual pre-agreement average. In terms of the notation introduced in Section 5.3, this results in profits $\pi_i(\{i\})$ for the focal doctor who leaves the network unilaterally. We calculate deviation profits relative to $\pi_i(\emptyset)$, i.e., the profits in the case were no doctor leaves the network.

Figure 10 plots the relative profits for every doctor from unilaterally becoming out-of-network $\frac{\pi_i(i)-\pi(\emptyset)}{\pi(\emptyset)}$, as well as the relative profits from jointly becoming out-of-network as in the Association $\frac{\pi_i(\mathcal{S})-\pi(\emptyset)}{\pi(\emptyset)}$.

[INSERT FIGURE 10 HERE]

The figure shows that most physicians would have had higher profits from leaving the network and setting their optimal price (assuming that all other doctors would have stayed in the network) relative to staying in-network. However, for several physicians the gains from deviating were small. In fact, one third of physicians would have increased their profits by less than 10 percent by deviating unilaterally. The figure also shows that the Association provides much higher profits than the in-network *status-quo* or than unilateral deviations for every doctor. On average, joint deviation produced profits 400 percent higher than the status quo, and 140 percent higher profits than unilateral deviation.

Two main insights emerge from this result. The fact that doctors did not leave the network unilaterally is consistent with the case documents showing that the physicians' goal was to reach ultimately better terms with the insurers, and that coordinated action was used to increase their bargaining position. In turn, coordinated action raised their outside option to the point where the negotiation broke down. The facts that unilateral deviations did not occur despite being incentive compatible, and that coordination entailed higher profits suggest that communication played an important role in the breakdown of negotiations.

6 Conclusion

Collusion among a large number of agents through a trade association is a common phenomenon. In this paper we study the collusive strategies of a trade association of physicians in the context of a vertical negotiation with payers. The case closely mirror several cases challenged by the FTC, in spite of recent calls by practitioners that put into question the prosecution of agreements that seek to counteract large imbalances of power in negotiations. 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.

The counterfactual exercises allow us to understand the stability and the coordination role of the association. First, we find that unilateral deviation profits from joint contract termination are much lower than the collusive profits. This result indicates that the Association's collusive strategy was incentive compatible. Second, we find that physicians would have benefited from leaving insurers unilaterally, but that they benefited much more from leaving insurers jointly. The fact that, in reality, physicians left insurers jointly is consistent with the physicians wanting to reach ultimately better terms with the insurers and with coordinated action being used to increase their bargaining position *vis-à-vis* the insurers. These results suggest that coordination played a key role in the price increases of 2012.

Our work has antitrust implications. Some antitrust practitioners have argued that small downstream firms should be allowed to jointly negotiate with larger upstream firms in order to "countervail market power," as most of the surplus is extracted by the large firm. We show that in the case we study joint negotiation caused large price hikes to consumers. This finding provides empirical support to the claim that such joint negotiations should be disallowed when the policymaker's goal is to maximize consumer surplus.

In addition, the Gynecologists' Association did not manage to raise prices much above the competitive outcome despite sustained coordination. 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.

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 the Association increased the profits of doctors' outside option in the bargaining game so that there was no longer surplus created by the agreement with insurers. This fact 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 a promising area for future research.

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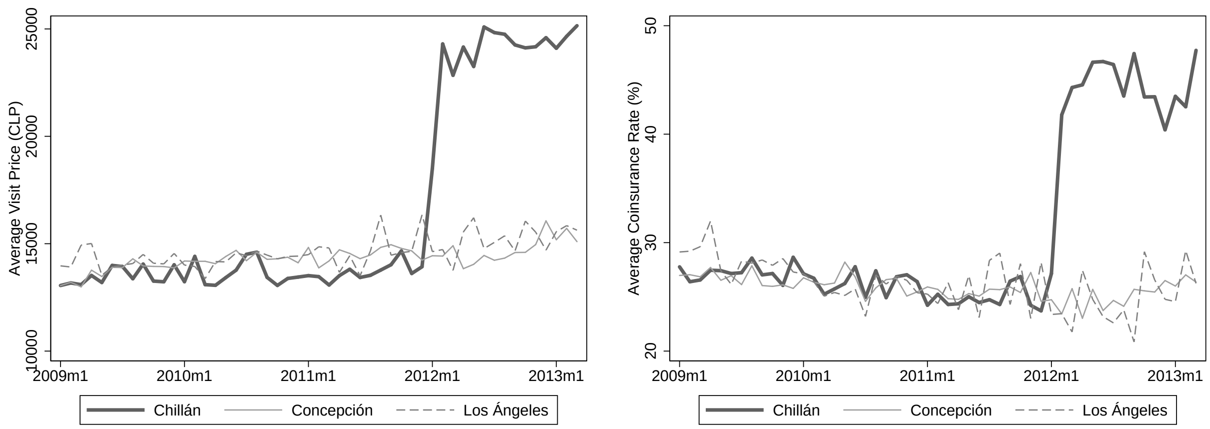
Figures

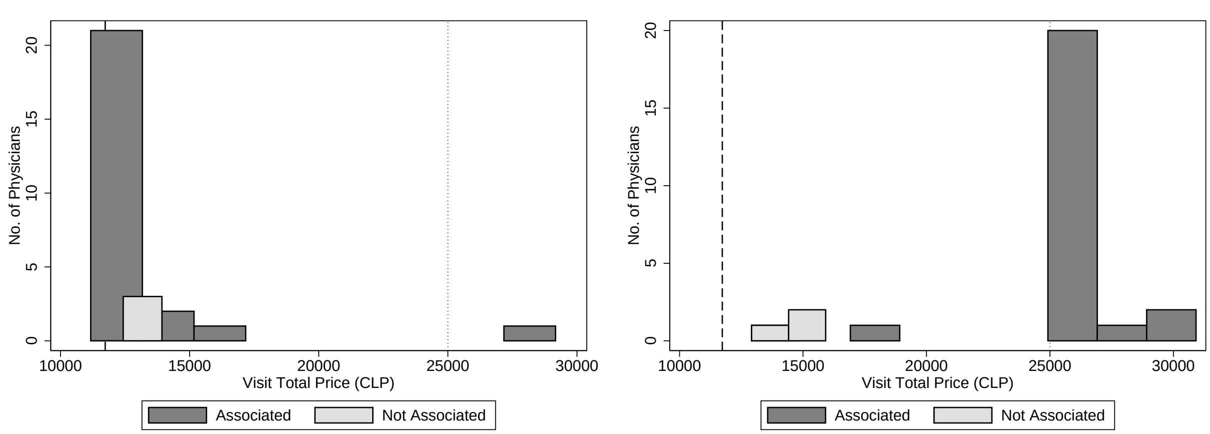
Figure 1: Prices and Coinsurance rates over time



(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).

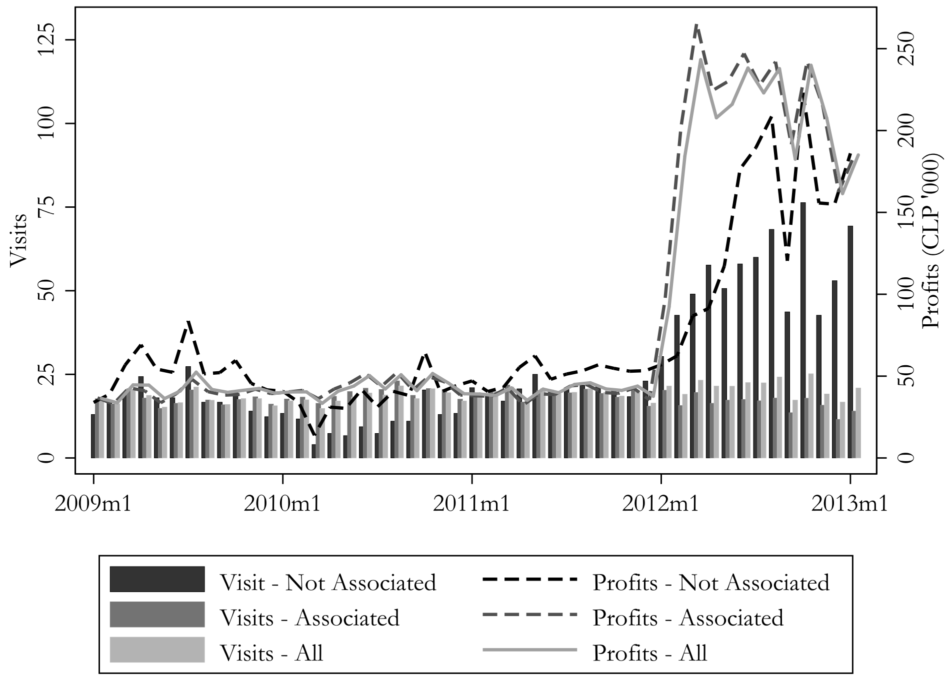
Figure 2: Histogram of Median Prices



(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.

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



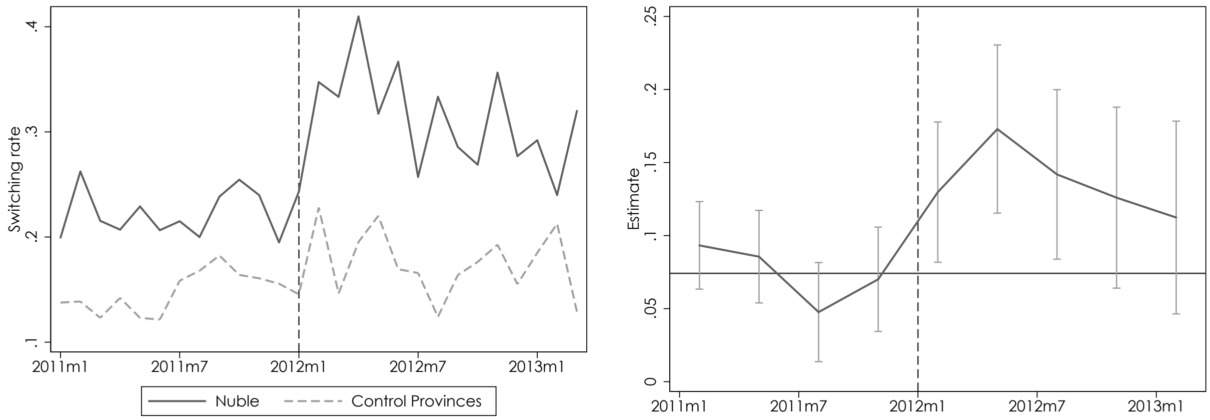
Note: The Figure shows the average number of visits and profits (using the FONASA payment as the cost) for all doctors combined and for groups of associated and not associated physicians.

Figure 4: Insurer Expenses and Out-of-Pocket Expenses

Note: The Figure shows patients' and insurers' aggregated expenses of visits.

The sum of the two lines represent physicians' revenues.

Figure 5: Physician switching rates in Ñuble and surrounding provinces

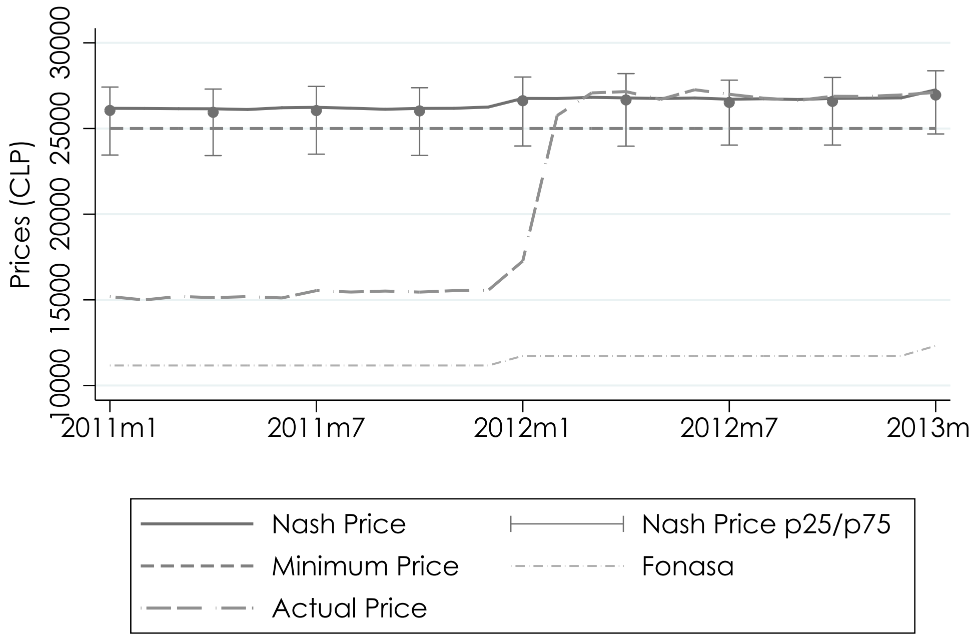


(a) Switching Rates

(b) Difference

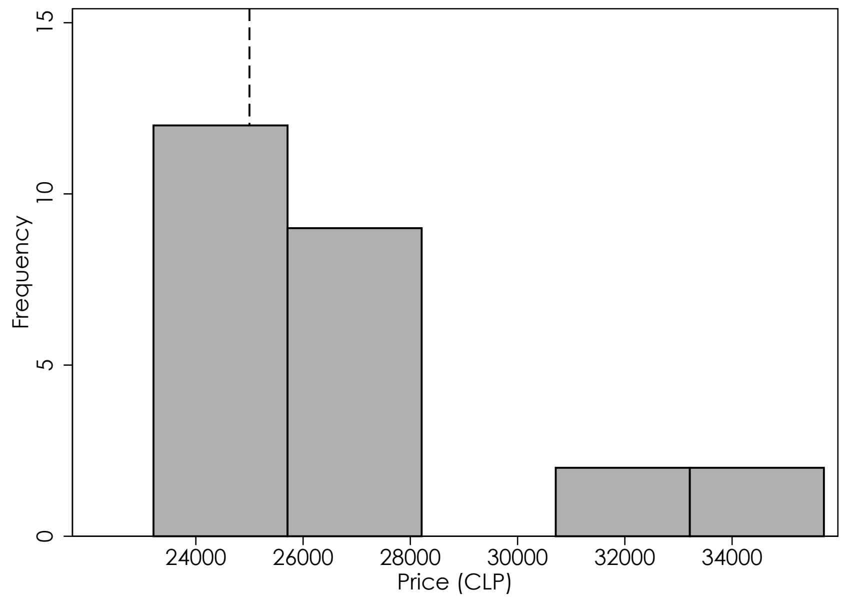
Note: Panel (a) shows the average switching rates across doctors for OBGyN visits in the province of Ñuble (where collusion occurred) and in surrounding provinces (Biobío and Concepción). Panel (b) shows the quarterly difference in switching rates between the two groups and their 95 percent confidence intervals. The vertical line marks the implementation date of the collusive measures. The horizontal line in Panel (b) shows the estimates average in 2011.

Figure 6: Simulated and Actual Prices



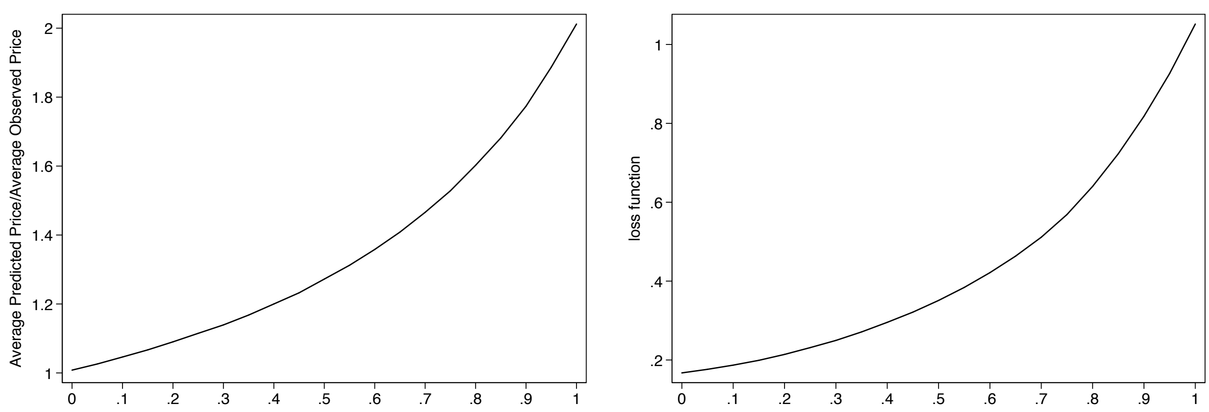
Note: The Figure shows the Nash prices for out-of-network coinsurance rates, the average actual price, and the FONASA rate. The dashed line shows the minimum price set by the Association. In addition, the vertical lines show the 25 and 75 percentiles of the out-of-network Nash prices, and the dot indicates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

Figure 7: 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.

Figure 8: Equilibrium Prices and Fit as a function of κ

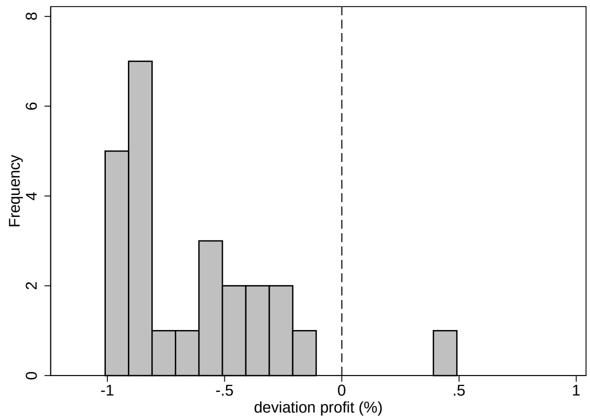


(a) Predicted relative to observed

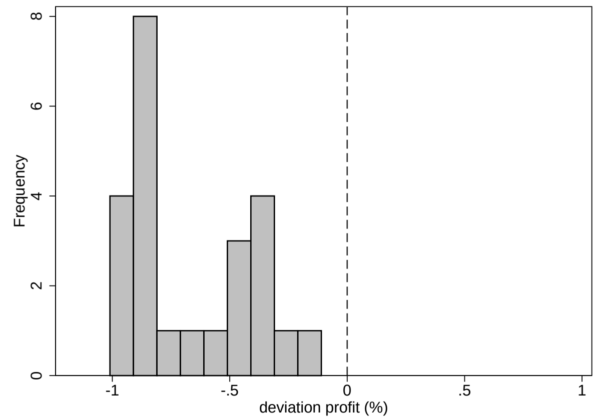
(b) Absolute error (loss function)

Note: Panel (a) shows the ratio between the sales-weighted average price predicted by the supply model and the sales-weighted average observed price in the post-association period. Panel (b) shows the "loss function" $f(\kappa) \equiv \frac{1}{T}\sum_{t\geq t_0} \frac{\frac{1}{N}\sum_i |1-\frac{p^N_{it}(\kappa)}{P_{it}}| q_{ijt}(\kappa)}{\sum_i q_{ijt}(\kappa)}$,

Figure 9: Deviation Profits and Association Stability



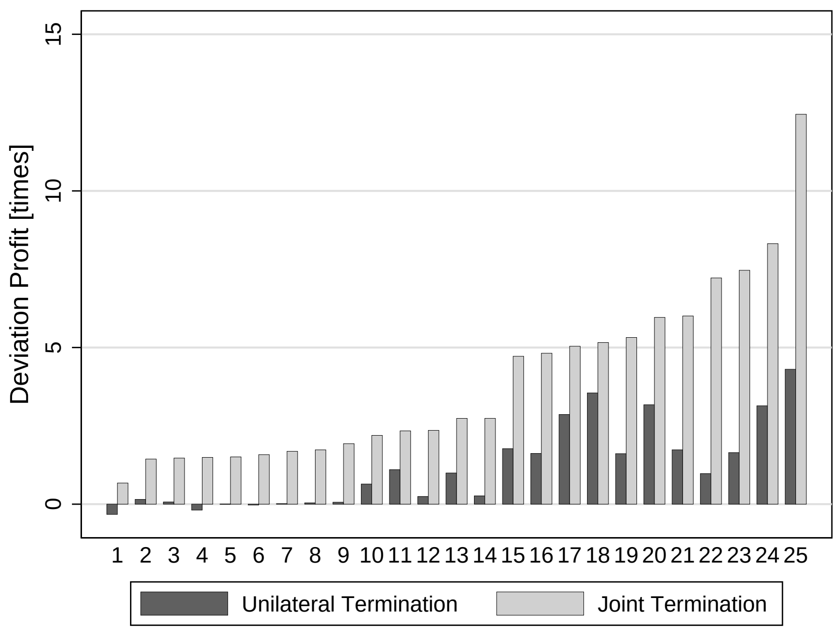
(a) Step 1



(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.

Figure 10: Deviation Profits as a Share of In-network Profits



Note: The Figure plots the gains from leaving the network under the assumption that all doctors of the Association leave the network (Joint Termination) and under the assumption that only the focal doctor becomes out-of-network (Unilateral Termination). Deviation profits are calculated as the difference in profits between each scenario and the profits under the situation that all doctors stay in-network.

Tables

Table 1: Physician switching rates

	(1)	(2)
$(After_t) \times T_l$	0.07***	0.07***
	(0.02)	(0.02)
$T_l = 1$	0.08***	0.08***
	(0.01)	(0.01)
$After_t$	0.03***	
	(0.01)	
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

Table 2: Demand Estimates

	(1)	(2)	(3)	(4)
	Logit OLS	Logit IV	NL OLS	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)
η_{pre}	0.02	-0.60	-0.29	-0.65
η_{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. * $p<0.1$, ** $p<0.05$, *** $p<0.01$

Appendix

A1 Differences-in-Differences Estimates for Prices and Coinsurance Rates

Table A1 shows the difference-in-differences results that we show graphically in Figure 1. The results indicate that as a result of the Association formation prices in Chillán increased CLP 8,160 and coinsurance rates increased 16 percentage points.

Table A1: Changes in List Price and Coinsurance Rate in Chillán and Neighboring Cities

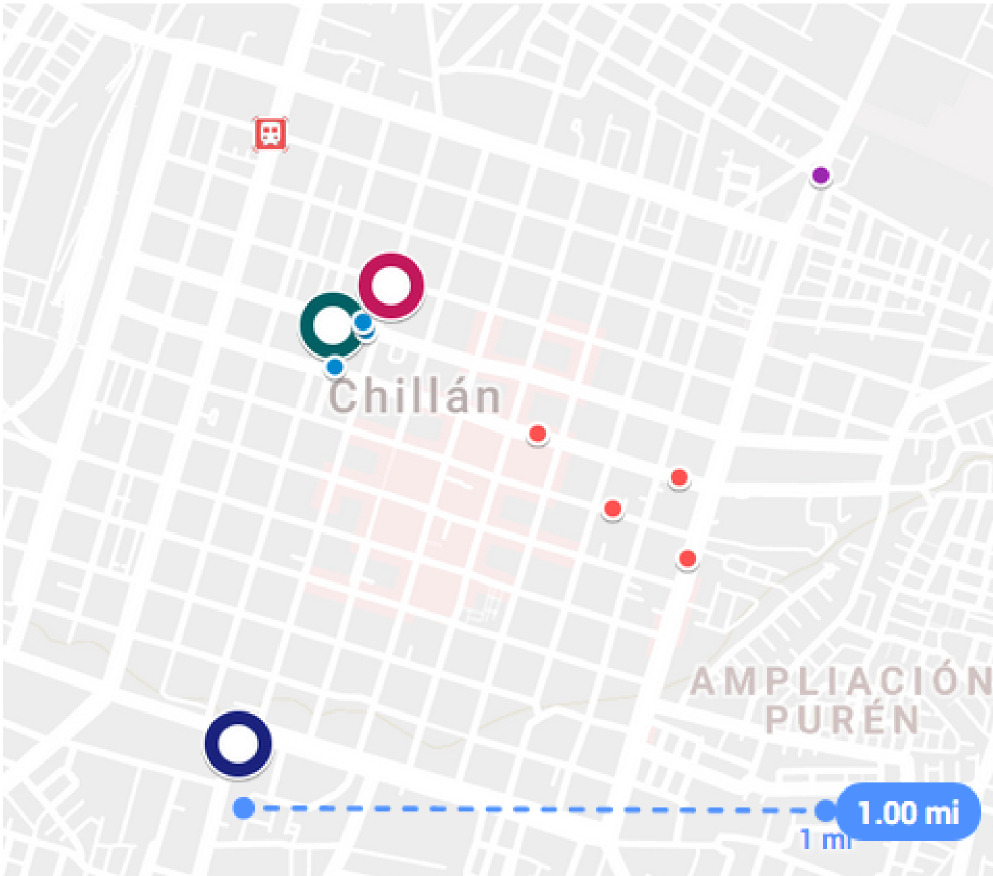
	(1)	(2)
	List Price	Coinsurance Rate
Chillán	-815.57***	-1.11***
	(145.11)	(0.30)
Post	4204.62***	3.29**
	(635.09)	(1.34)
Chillán x Post	8159.38***	15.89***
	(358.25)	(1.02)
Baseline	13731.46	26.18
N	5,850	5,813
R^2	0.21	0.12

Heteroscedasticity robust standard errors in parentheses. * $p<0.10$, ** $p<0.05$, *** $p<0.01$

A2 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. Physicians who did not join the Association are in the light-blue nest and in the orange nest.

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.

A3 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.* (2014). 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 A2: Nested Logit Estimates–Zero Market Shares Correction

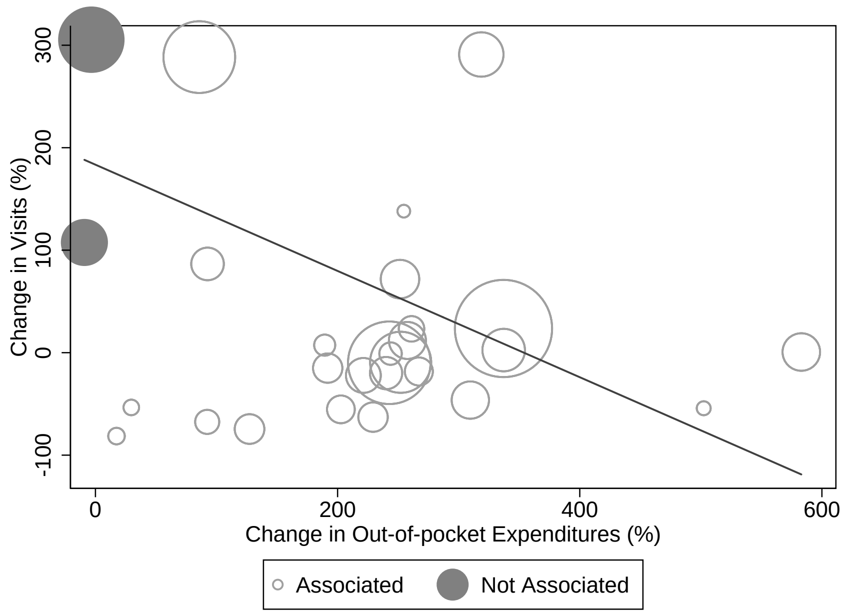
	(1)	(2)	(3)
$p_{ijt}(1 - c_{ijt})$	-0.015	-0.028***	-0.050***
	(0.011)	(0.008)	(0.018)
ln within-share	1.066***	0.850***	0.735***
	(0.148)	(0.108)	(0.152)
Correction	No Zeros	Ones	Dirichlet
η_{pre}	0.62	-0.61	-0.65
η_{post}	1.63	-1.58	-1.68
Observations	4870	6620	6620
AR F-stat	52.04	73.72	39.42
AP F-stat 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. * $p<0.1$, ** $p<0.05$, *** $p<0.01$

A4 Market Share Changes

Figure A2 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 in our sample increased their visits by 206 percent on average, while the median colluding doctor decreased their visits by 12 percent.

Figure A2: 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.

A5 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 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 (1) 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.

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

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. [47]

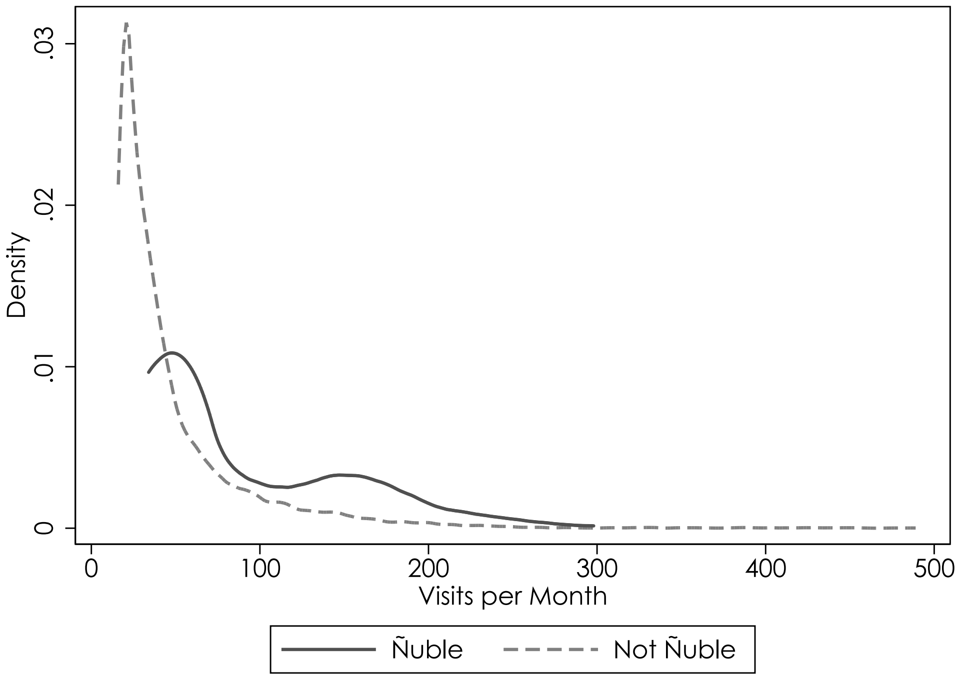
A6 Capacity Constraints

In this section we provide evidence that indicates no observable binding capacity constraints of patients visits *in the private sector*. First, we plot the left-truncated distribution of Isapre visits per month of all ob/gyns in Ñuble and in other neighboring provinces in Figure A3. Specifically, we plot the distribution of monthly visits between the 75 and the 99.9 percentile of the physician visits distribution (we right-censor the distribution due to possible measurement error). Although the figure shows substantial heterogeneity, 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

[47] 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

capacity constraints are not binding in Ñuble.

Figure A3: Truncated Distribution of Monthly Visits



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.

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 of these numbers, the average ob/gyn visit in the US lasts roughly 20 minutes.[48] Hence, a doctor working an 8-hour shift, 5 days a week, is able to see up to 480 patients a month. If the doctor devotes one full day only to procedures, then she is able to see up to 384 patients a month.[49] These numbers entail that 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.

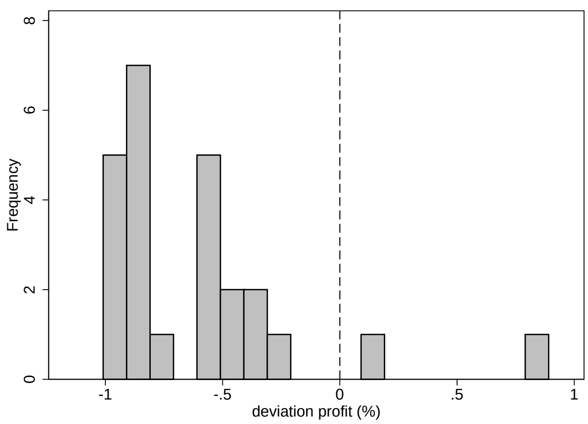
[48]See "Profile of Ob-Gyn Practice 1991-2003", *American College of Obstetricians and Gynecologists*, `https://www.acog.org/~/media/Departments/Practice/ProfileofOb-gynPractice1991-2003.pdf`. Accessed Oct 25, 2018.

[49]In the pre-Association period there was 1 procedure for every 23 visits on average in all the cities in our sample. If a procedure takes one hour on average, this means that ob/gyns devote half a day to procedures.

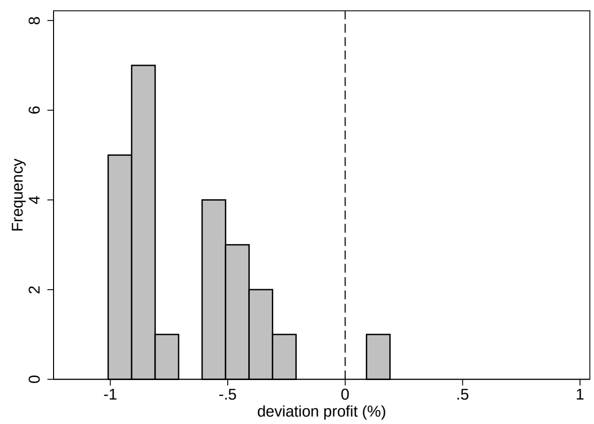
A7 Incentives to Join

Figure A4 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 9 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.

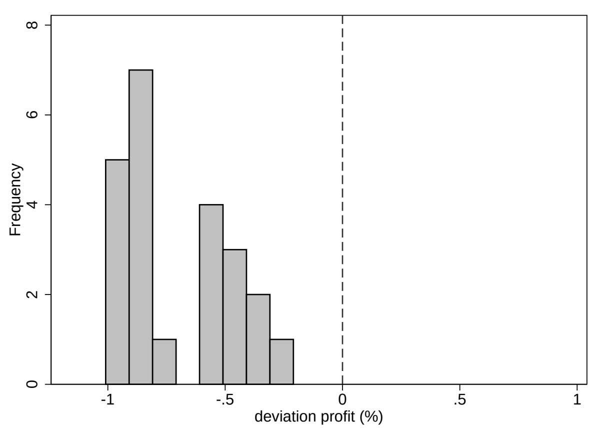
Figure A4: Deviation Profits and Association Stability



(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.