Buyer Alliances in Vertically Related Markets

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Abstract

Alliances of buyers to negotiate input prices with suppliers are commonplace. Using pre- and post-alliances data on bottled water purchases, I develop a structural model of bilateral oligopoly to estimate the effects of three alliances formed by retailers on their bargaining power vis-à-vis manufacturers and retail prices paid by consumers. The results provide evidence of a countervailing buyer power effect that reduces retail prices by more than 7%. Exploring determinants of buyer power, I find that changes in the bargaining ability of retailers play an important role in the countervailing force of the alliances, which otherwise would not have been profitable.

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JEL classification: C78, D43, L11, L13, L14, L41, L81.

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1 Introduction

The formation of alliances by groups of economic agents to bargain with other agents is a widespread phenomenon. Typical examples include labor unions which negotiate wages with employers on behalf of workers, buyer alliances formed by retailers to negotiate wholesale prices with food manufacturers (Colen et al., 2020), group purchasing organizations through which hospitals join forces to bargain with their suppliers (Burns and Lee, 2008).[1] How collective bargaining and market concentration affect negotiation outcomes has long been a subject of economic inquiry (Segal, 2003). Although conventional wisdom suggests that size and group membership confer a bargaining advantage (Galbraith, 1952), theory provides ambiguous predictions and empirical evidence remains sparse.[2]

Leveraging a unique case on the French food retail sector in 2014, this article attempts to provide a comprehensive analysis of the effects of three buyer alliances formed by retailers on their bargaining with manufacturers and retail prices paid by consumers. For tractability, I focus on the bottled water sector, which features a classic example of bilateral oligopoly and includes some of the top-selling items in the French supermarket industry.[3] Using household-level scanner data on bottled water purchases for the years 2013 (pre-alliances periods) and 2015 (post-alliances periods), I take advantage of the quasi-experimental variation created by the formation of buyer alliances to explore determinants of buyer power and analyze their effects on market outcomes.

My empirical approach is outlined as follows. I start with a descriptive analysis that provides evidence of a substantial decline in the retail prices of national brands sold by retailers involved in an alliance following the formation of buyer alliances. To explore the mechanisms underlying this retail price drop and analyze the effects on the surplus division in the vertical chain, I use a structural model of demand and supply. The demand-side consists of a standard discrete choice model that incorporates both observed and unobserved heterogeneity in consumer preferences to accommodate rich substitution and curvature patterns (e.g., Berry, Levinsohn and Pakes, 2004). The supply-side extends Bonnet, Bouamra-Mechemache

[1] Other real-world examples can be found in the pharmaceutical industry (e.g., pharmacy benefit managers negotiate drug prices for health insurers and employers), the market for academic journals (Jeon and Menicucci, 2017), the U.S. cable television industry (Chipty and Snyder, 1999), the U.S. retail hardware market, or the aircraft sector (Dana, 2012).

[2] For example, Chipty and Snyder (1999) provide ambiguous theoretical results and find no empirical support for the claim that horizontal mergers confer a bargaining advantage in the cable television industry. DiNardo and Lee (2004) obtain similar findings regarding unionization.

[3] According to NielsenIQ, the best-selling item in supermarkets is the brand of bottled water Cristaline, where nearly one-third of households have purchased it at least once in 2018. Additionally, 4 brands of bottled water are among the top 10 items generating the highest revenue in the supermarket industry.

and Molina (forthcoming) by developing an empirical framework suitable for the analysis of buyer alliances in bilateral oligopolies. The framework includes a two-stage game in which manufacturers and retailers engage in bilateral negotiations to determine wholesale prices of products and where retailers subsequently compete in retail prices on the downstream market. Given the presence of contracting externalities, I use the "Nash-in-Nash" bargaining solution (Horn and Wolinsky, 1988*a*) as a surplus division rule in the vertical chain.

Lacking data on negotiated wholesale prices, I rely on estimates of consumer demand as well as on the set of equations characterizing necessary conditions for a Nash equilibrium in retail prices to recover the marginal costs of retailers before and after the formation of buyer alliances (e.g., Villas-Boas, 2007). To analyze the determinants of these inferred marginal costs, I specify a retail marginal cost function in terms of two additive components. The first component corresponds to the marginal cost of production and distribution for bottled water products, which is constant in quantity and includes a structural error term reflecting unobserved cost factors as in Gowrisankaran, Nevo and Town (2015). The second component is the price-cost margins of manufacturers resulting from the negotiation process with retailers and whose (closed-form) expression derives from the first-order conditions of the "Nash-in-Nash" bargaining solution. A key innovation of my approach is to exploit the variation in the (inferred) marginal costs of retailers caused by the alliances to estimate changes in three determinants of bargaining power: (i) the status quo position of manufacturers and retailers in each bilateral negotiation, (ii) their costs from making price concessions during negotiations, and (iii) their bargaining ability.

Given that the observed retail price change following the alliances can be rationalized by unobserved shocks, I place restrictions on demand and cost unobservables that are similar in spirit to Miller and Weinberg (2017). In particular, my identification assumptions rely on the presence of products remaining outside the scope of the alliances to control for common changes in demand and cost unobservables before versus after the formation of buyer alliances. Then, conditional on a set of control variables, I interpret the formation of buyer alliances as an exogenous shifter of the bargaining environment in the vertical chain to form moment conditions and estimate structural parameters.

The results show that the share of industry profit captured by retailers increases from 65.19% in the pre-alliances periods to 87.25% in the post-alliances periods, indicating that retailers have gained bargaining power vis-à-vis manufacturers. To quantify the effects attributable to the formation of the alliances, I leverage the model and parameter estimates to simulate equilibrium market outcomes absent buyer alliances. I find that the observed post-alliances retail prices of national brands sold by retailers involved in an alliance are, on average, 8.15% lower than they would have been absent buyer alliances. I also find that the (quantity-weighted) price-cost margins of manufacturers and the industry profit are respectively 71.22% and 10.06% lower relative to the counterfactual scenario without the alliances. This result provides empirical evidence of a countervailing buyer power effect that generates a substantial drop in wholesale and retail prices to the detriment of manufacturers and the bottled water industry. In two other counterfactual experiments, I investigate the determinants underlying buyer power in the vertical chain. I find that retailers' asymmetries and changes in the relative bargaining ability of retailers vis-à-vis manufacturers play an important role in the countervailing force and the profitability of buyer alliances. Specifically, when there is no increase in retailers' bargaining ability, the results indicate that the largest retailer in each buyer alliance is worse off bargaining as an alliance member than bargaining alone. This finding is reminiscent of the joint-bargaining paradox first observed by Harsanyi (1977). I show that this paradox stems from a bargaining force previously unexplored in empirical work: the nondiscrimination effect of buyer alliances. First, this effect impacts the willingness of manufacturers and retailers to accept price concessions in negotiations. Second, it reduces the competitive advantage of the large retailer over its downstream rival because both benefit from the same purchasing conditions.

The present article contributes to the extensive literature on buyer power that, dating back to Galbraith (1952) and its concept of countervailing power, analyzes the potential for large buyers to secure lower input prices (see Snyder, 2008, for a comprehensive survey). Earlier theoretical works on buyer alliances and unionization have identified that the concavity of the gains that a manufacturer gets from reaching an agreement is a key determinant of countervailing power.[4,5] A recent stream of the literature has also found that the ability of a buyer alliance to coordinate the purchasing policy of its members is an important source of countervailing buyer power.[6] My contribution to this line of research is twofold. On the one hand, I highlight that a buyer alliance enhances the relative status quo position

[4]See Horn and Wolinsky (1988*b*) when workers are substitutes and Chipty and Snyder (1999) when costs are convex. Inderst and Montez (2019) have shown that this condition does not necessarily extend to settings with multiple manufacturers in which adjustments of trades are allowed upon bilateral disagreement.

[5]Theoretical analysis of downstream market concentration (e.g., retail merger) have also examined conditions for countervailing buyer power (Dobson and Waterson, 1997; Iozzi and Valletti, 2014; Gaudin, 2018). A critical difference with buyer alliances is that the reduction in the number of retailers at the downstream level puts upward pressure on retail prices, which tends to blur the analysis of pure countervailing power effects.

[6]For instance, an alliance can heighten upstream competition by pre-committing to reduce the number of manufacturers to deal with (e.g., Inderst and Shaffer, 2007; Dana, 2012; Chen and Li, 2013; Allain, Avignon and Chambolle, 2020). An alliance can also improve the status quo position of retailers by engaging in negotiations with multiple manufacturers (Chae and Heidhues, 2004*a*), or increase the outside option of retailers when dealing with a powerful manufacturer (Caprice and Rey, 2015).

of its members vis-à-vis manufacturers in the event of bilateral disagreement.[7] On the other hand, I show that the countervailing force of this status quo effect may be undermined by the nondiscrimination effect of a buyer alliance that negotiates similar trading terms for its members, resulting in ambiguous effects on buyer power.

This article also builds upon a recent body of empirical works that use the "Nash-in-Nash" bargaining solution to estimate models of buyer-seller relationships with contracting externalities (e.g., Draganska, Klapper and Villas-Boas, 2010; Crawford and Yurukoglu, 2012). Prior articles in this literature have primarily relied on ex-ante structural analysis to study the effects of market concentration on bargaining outcomes.[8] For instance, Gowrisankaran, Nevo and Town (2015) find that a hospital merger in Northern Virginia would have significantly raised prices despite the buyer power of insurers. Bonnet, Bouamra-Mechemache and Molina (forthcoming) highlight that a merger between two competing retailers may weaken their bargaining power vis-à-vis soft drink manufacturers, resulting in higher wholesale and retail prices. Ho and Lee (2017) find that insurer concentration (by removing an insurer) may lead to lower hospital prices and premiums when employers impose premium constraints through bargaining. Using an oligopoly model of bargaining, Grennan (2013) estimates the likely effects of a group purchasing organization negotiating with medical device manufacturers on behalf of multiple independent hospitals. He shows that the profitability of group purchasing ultimately depends on an increase in the bargaining ability of hospitals vis-à-vis manufacturers. My contribution to this literature is to provide an ex-post analysis of market concentration through buyer alliances. I extend Grennan's (2013) work to bilateral oligopolies and provide more direct evidence that shifts in bargaining abilities constitute an important source of profitability for buyer alliances.[9] I also shed light on a countervailing buyer power effect which, unlike Ho and Lee (2017), benefits both retailers and consumers and emerges without any constraints on retail price setting. Consistent with Galbraith's (1952) view, these findings offer new empirical evidence on the underlying determinants of buyer power and their effects on welfare which, since the substantial rise of large retailers these last decades (Hortaçsu and Syverson, 2015; Smith and Ocampo, forthcoming), have become a central issue for competition authorities.[10]

[7]This bargaining effect can be related to the curvature of the manufacturer's gains from reaching an agreement, which is concave when retailers are substitutes (e.g., Horn and Wolinsky, 1988*b*).

[8]Another strand of the literature has studied the effects of mergers (e.g., Lewis and Pflum, 2017; Craig, Grennan and Swanson, 2021) or buyer alliances (Dubois, Lefouili and Straub, 2021) on negotiated prices using reduced-form analysis. Though insightful, this approach provides limited guidance on the mechanisms underlying buyer power and their implications for welfare.

[9]This finding can also be related to Lewis and Pflum (2015) who estimate that greater bargaining ability enables hospital systems to negotiate higher prices with insurers compared to individual hospitals.

[10]As emphasized in Carlton and Israel (2011), the proper antitrust treatment of conduct and market struc-

The remainder of this article is structured as follows. Section 2 sheds light on the main economic forces at play when two retailers form a buyer alliance. Section 3 introduces relevant features of the buyer alliances in the French food retail sector, the data used in the empirical analysis, and reduced-form evidence on the evolution of retail prices before and after the formation of buyer alliances. Section 4 describes the structural model of demand and supply, discusses identification and estimation, and presents the empirical results. Section 5 presents the counterfactual simulations and Section 6 concludes. Technical details and robustness checks are deferred to the Supplemental Material.

2 Theoretical insights

The effects generated by a buyer alliance can be appreciated in a simple setting of vertical relations as illustrated in Figure 1. A monopoly manufacturer, M, sells its brand to two symmetric retailers, R_1 and R_2, indexed by $r = 1, 2$, competing for consumers in a downstream market. Firms operate under constant returns to scale, incurring a unit cost of production and distribution normalized to 0 for simplicity.

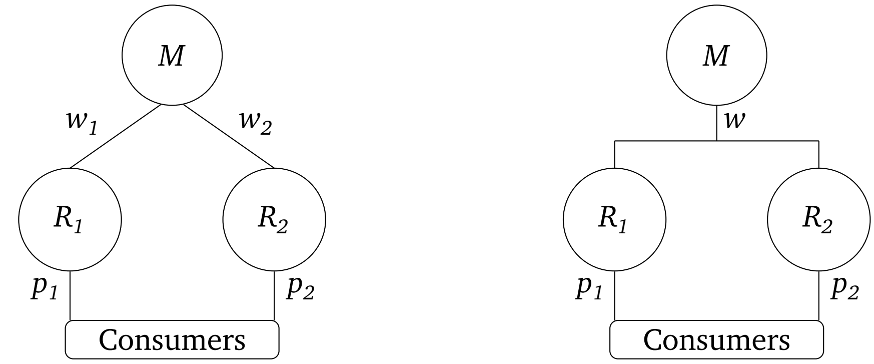
In the benchmark case depicted in Figure 1a, each retailer r simultaneously and secretly engages in a bilateral negotiation with M to determine its linear wholesale price w_r before competing for consumers by setting its retail price p_r. A modeling difficulty of the bargaining game is the prevalence of contracting externalities due to the downstream competition (that is, the surplus to be divided in one negotiation depends on the outcome of the other negotiation). To overcome this issue in a tractable way, the literature has leveraged the surplus sharing rule pioneered by Horn and Wolinsky (1988*a*) and commonly referred to as "Nash-in-Nash" bargaining solution. This concept considers that each wholesale price is determined according to the two-person Nash bargaining solution (Nash, 1950), taking the wholesale prices from other bargains as given.[11] Applied to the benchmark setting where retailers face a monopoly manufacturer, the "Nash-in-Nash" bargaining solution implies that the status quo payoff of retailer r in its negotiation with M is 0. However, purchases by the other retailer confer a positive status quo payoff to M in its negotiation with retailer r. As a result, M has a stronger bargaining position than the retailers in each bilateral bargain.

In the situation depicted in Figure 1b, retailers join forces by negotiating a common

ture changes that create or enhance buyer power remains an unsettled question. For example, the Federal Trade Commission and the Department of Justice have adopted conflicting views on the treatment of buyer power in recent merger reviews (Hemphill and Rose, 2018).

[11] The terminology "Nash-in-Nash" has been coined by Collard-Wexler, Gowrisankaran and Lee (2019) who highlight that this bargaining solution can be interpreted as a Nash equilibrium in wholesale prices negotiated by pairs of firms according to the Nash's axiomatic theory of bargaining. See Section 4.2 for further details.

Figure 1: Vertical market structure



(a) Benchmark: No alliance (b) Buyer alliance

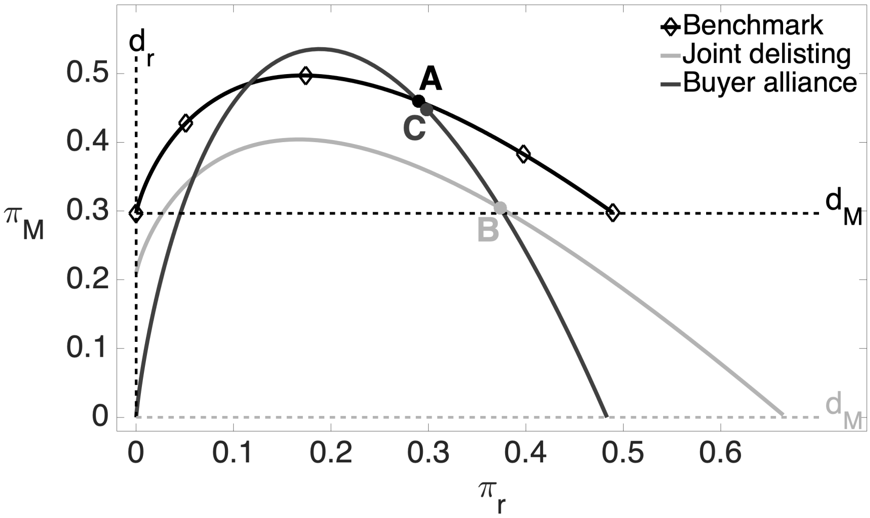
wholesale price w through a buyer alliance that aims at maximizing their joint profit. Hence, only one bilateral negotiation remains affecting (i) the relative status quo position of firms, (ii) their costs of making price concessions, and (iii) their bargaining ability. In what follows, I provide a non-technical discussion of each bargaining effect using Figures 2 and 3. I refer to Section S1 of the Supplemental Material for a more formal analysis.

Status quo effect (joint delisting decision). The status quo effect of a buyer alliance can be grasped by considering an intermediate case in which retailers join forces through a buyer alliance but keep negotiating their wholesale prices separately and secretly. By coordinating the purchasing policy of its members, I assume that the alliance endows each retailer a veto power in the spirit of Caprice and Rey (2015). Thus, in the event of bilateral disagreement with one retailer, all retailers in the alliance jointly delist M's brand from their shelves, which deteriorates the status quo payoff of M and undermines its bargaining power vis-à-vis retailers.[12]

Figure 2 illustrates this effect by characterizing different bargaining situations between M and retailer r, taking as given the wholesale price negotiated with the other retailer. The x-axis measures retailer r's profit (π_r) and the y-axis M's profit (π_M). Starting from the benchmark case, the status quo payoff of M (d_M) and retailer r (d_r) are respectively depicted by the horizontal and vertical black dotted lines. The black curve with diamond markers represents the set of all feasible agreement points (bargaining frontier), where the symmetric Nash bargaining solution is located at point A (with bargaining weights of firms equal to 0.5). Note that the concave shape of the bargaining frontier stems from the fact that the negotiated price w_r affects both the allocation of surplus between firms (π_M and π_r)

[12] Note that this effect is absent in Caprice and Rey (2015) who instead focus on the polar case in which M makes take-it-or-leave-it offers, implying that its status quo position is irrelevant to the bargaining outcome.

Figure 2: Effects on the bargaining frontier and the Nash solution



Notes: This figure depicts different bargaining situations between M and retailer r. It is drawn under a logit demand system, two symmetric retailers, and zero marginal cost of production and distribution (see Section S2 of the Supplemental Material for further details on the generating process). The x-axis measures retailer r's profit (π_r) and the y-axis measures M's profit (π_M). The horizontal and vertical dotted lines show respectively the status quo payoff of M (d_M) and retailer r (d_r). The black curve with diamond markers, the grey, and the black curve without markers represent the set of all feasible agreement points (bargaining frontier) under the benchmark case, the joint delisting decision case, and the buyer alliance case respectively. The point A, B, and C locate the symmetric Nash bargaining solution on each bargaining frontier.

and the total surplus to be divided ($\pi_M + \pi_r$).[13] As previously described, the buyer alliance decreases M's status quo payoff from 0.3 to 0, which is illustrated by the downward shift in d_M (grey dotted line). Under this joint delisting decision scenario, the bargaining frontier is represented by the grey curve and the symmetric Nash bargaining solution by point B. Hence, the status quo effect of the buyer alliance increases retailers' profits at the expense of M.

It is worth mentioning that the countervailing force of this status quo effect can be more broadly related to a result first identified in the wage bargaining literature (e.g., Horn and Wolinsky, 1988*b*). In particular, collective bargaining is beneficial when the gains perceived by the employer (manufacturer) from its bilateral agreements are concave, which arises when workers (retailers) are substitutes. In settings of vertical relations with linear contracting, I show that collective bargaining generates another force previously unexplored in the literature: the nondiscrimination effect.

[13]This is a consequence of the well-known double marginalization (Spengler, 1950). As w_r increases, π_M increases and π_r decreases, but $\pi_M + \pi_r$ may either increase or decrease (for large values of w_r, π_M may decrease as well). Under efficient contracting (e.g., lump-sum payments), the shape of this bargaining frontier would instead be a straight line with slope -1 (see, e.g., Grennan, 2014).

Nondiscrimination effect. The buyer alliance described in Figure 1b allows retailers to obtain similar trading terms when purchasing M's brand. Hence, there is just one bilateral negotiation in which M cannot price-discriminate between retailers. As formally shown in Section S1 of the Supplemental Material, this nondiscrimination effect alters firms' relative costs of making price concessions in bargaining (see also O'Brien, 2014, in the context of input price discrimination).[14] The intuition is as follows. On the one hand, a retailer making a price concession to M no longer suffers from a competitive disadvantage in the downstream market due to a higher marginal cost because this concession is shared with its rival. On the other hand, M's concession cost increases because price concessions are given to both retailers at the same time. Consequently, the nondiscrimination effect of a buyer alliance implies that the costs of making price concessions are less (resp. more) painful for retailers (resp. M), which reinforces the bargaining power of M vis-à-vis retailers. This effect is illustrated in Figure 2 where the black curve represents the bargaining frontier under the buyer alliance. Compared to the grey curve in which wholesale price discrimination arises, the slope of the black curve is steeper. This indicates that it is easier (resp. harder) to transfer surplus from the retailers (resp. M) to M (resp. the retailers) through the wholesale price w.[15] The symmetric Nash bargaining solution located at point C shows that the division of surplus shifts to the benefit of M.

Bargaining ability effect. Some recent empirical works have highlighted that changes in market conditions are likely to affect the bargaining ability of firms, which may have important implications for predicting market outcomes (e.g., Grennan, 2013; Lewis and Pflum, 2015; Grennan and Swanson, 2020).[16] Although various interpretations have been advanced to explain asymmetries in the bargaining ability of firms,[17] the literature on n-person bargaining games can provide some insights into the effects of buyer alliances. Kalai (1977) introduced the asymmetric Nash bargaining solution on the ground that a bargainer representing the interest of multiple players should be treated more favorably by the Nash

[14]The cost of making a price concession can be understood as the marginal effect of agreeing upon a higher (resp. lower) wholesale price on the retailer's profit (resp. M's profit).

[15]In other words, M has to bear more losses to increase retailer r's profit under the black curve than the grey curve.

[16]For instance, Grennan and Swanson (2020) provide evidence that heterogeneity in bargaining abilities can reflect asymmetric information. This underscores the importance of accounting for changes in the bargaining ability of firms when analyzing changes in market conditions that affect transparency.

[17]For instance, Roth (1979) suggests that the presence of asymmetries in firms' bargaining ability can be based on some information or other factors that are "outside" the model. Binmore, Rubinstein and Wolinsky (1986) argue that this can reflect asymmetries in the probability that each firm is selected to make a proposal at each bargaining period, asymmetries in the patience of firms, or asymmetries in firms' beliefs about the likelihood that a bargaining breakdown occurs due to an exogenous event.

solution (see also Roth, 1979).[18] By interpreting the situation depicted in Figure 1b as a 3-person bargaining game where the alliance negotiates on behalf of both retailers, Kalai (1977) offers support for a change in bargaining abilities to the benefit of retailers. In Figure 2, this effect would simply shift point C to the right along the black curve, thereby increasing the profit of retailers to the detriment of M. Although Kalai's (1977) theory provides an axiomatic foundation to the bargaining ability effect of buyer alliances, I acknowledge that a more formal grounding based on noncooperative bargaining games is beyond the scope of this article. Bearing these considerations in mind, I develop an empirical approach to infer changes in bargaining abilities while remaining fairly agnostic about the precise mechanisms at play.[19]

Retailers' asymmetries. Additional insights can be obtained from Figure 3 which presents the profitability of forming an alliance according to the bargaining weight of retailer 1 (x-axis) and retailer 2 (y-axis) vis-à-vis M. For the sake of simplicity, I consider that the bargaining weight of the alliance is equal to the highest bargaining weight of both retailers (i.e., the bargaining ability effect of the alliance only benefits the weakest retailer). The figure shows that asymmetries among retailers make the formation of an alliance less likely.[20] In particular, a powerful retailer that negotiates low wholesale prices has less incentive to form an alliance when its rival is a weak bargainer.[21] The insight underlying this instability result is that the (potential) gain of bargaining power from forming an alliance must compensate for the losses of having a more competitive downstream rival. By making the purchasing conditions more uniform across retailers, the profitability of a buyer alliance formed by two asymmetric retailers depends ultimately on a bargaining ability effect that benefits not only the weakest but also the strongest retailer. Interestingly, this finding echoes another result in the wage bargaining literature suggesting that unions among homogeneous workers are easier to form (Jun, 1989).

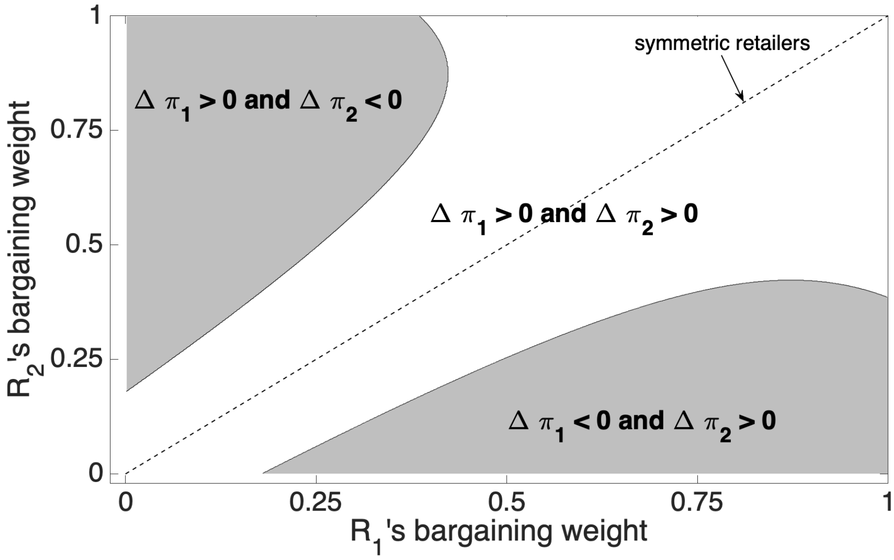
[18] A reformulation of this theory using the notion of multiple "right to talk" (or "right to make a proposal") has been suggested in the literature (e.g., Chae and Heidhues, 2004*b*). More specifically, in a random-proposer bargaining protocol where each firm has an exogenous probability of being selected to become a proposer, forming a coalition of multiple firms increases the chance of making proposals, which improves the bargaining power of coalition members vis-à-vis other trading partners.

[19] Other approaches impose further structure by parameterizing the bargaining ability in terms of observed firm and market characteristics (e.g., Lewis and Pflum, 2015; Leong et al., 2022).

[20] When a retailer's bargaining weight gets closer to 1 (i.e., it makes a take-it-or-leave-it offer to M), grey areas in which the alliance is unprofitable tend to shrink because the nondiscrimination effect vanishes. Despite this effect, a buyer alliance need not be profitable as it improves the competitiveness of the rival retailer in the downstream market.

[21] Although side payments could in principle address this issue, they are likely to violate antitrust law and be difficult to implement in practice (see, e.g., the discussion on page 64 of Inderst and Shaffer, 2007).

Figure 3: Profitability of a buyer alliance



Notes: This figure depicts changes in the profit of retailers following the formation of a buyer alliance ($\Delta\pi_r$). It is drawn under a setting similar to Figure 2, except that retailers may have asymmetric bargaining weights vis-à-vis M. Bargaining weights are from a two-dimensional grid where the x-coordinates and y-coordinates vary between 0 and 0.999 with an increment of 0.001. For every point in the grid, I solve the system (S10) in Section S2 of the Supplement Material to recover the equilibrium retail prices and profits of retailers. When retailers have asymmetric bargaining weights, I assume that the bargaining weight of the alliance equals the highest bargaining weight of both retailers.

Summary. Shedding light on the main bargaining forces at play, I show that a buyer alliance of two competing retailers generates ambiguous effects on buyer power. Although the status quo effect and the bargaining ability effect are likely to reinforce the bargaining power of retailers vis-à-vis manufacturers, the nondiscrimination effect reduces their strength. The profitability of forming a buyer alliance is also questionable as it undermines the competitive advantage of a retailer vis-à-vis its downstream rivals. In what follows, I analyze the formation of buyer alliances from an empirical perspective using insights developed in this section to guide my approach and interpret the economics behind my estimates.

3 Industry background, data and descriptive analysis

My empirical application focuses on the formation of three buyer alliances in the French food retail sector in 2014. This section introduces the relevant industry background, the alliances, the data, and a descriptive analysis of changes in retail prices following the year of the alliances.

3.1 Industry background

Seven large retailers (Carrefour, Cora, Groupe Auchan, Système U, ITM Entreprises, Groupe Casino, and Leclerc) as well as hard discounters (e.g., Lidl, Aldi) compete in the French food retail sector to attract consumers to their stores. Every year, from November to February, each retailer engages in annual negotiations with food manufacturers to determine wholesale prices of products.[22] In 2008, the "Loi de Modernisation Economique" allowed retailers to obtain discriminatory wholesale prices to intensify competition and decrease retail prices paid by consumers (Allain, Chambolle and Turolla, 2022). In 2014, six of the large retailers joined forces with one of its rivals to negotiate common wholesale prices, giving rise to the formation of three buyer alliances: (i) Carrefour and Cora, (ii) Système U and Groupe Auchan, and (iii) ITM Enterprises and Groupe Casino.[23] The scope covered by these alliances was limited to the wholesale price negotiations of products sold under national brands by large manufacturers only, thereby excluding private labels (store brands), fresh products (e.g., fruit and vegetables), and products supplied by small and medium enterprises. In collaboration with the French Senate, the Minister for the Economy consulted the French competition authority which issued an opinion on the potential anticompetitive effects of such operations (see Autorité de la concurrence, 2015). Recognizing that complex economic forces were at play, the authority did not challenge the alliances but advocated for more scrutiny by imposing a prior notification for the formation of any new buyer alliance in the food retail sector.

For tractability motives, my analysis focuses on the bottled water industry which constitutes an attractive laboratory for studying buyer alliances for at least three reasons. First, this industry is characterized by a bilateral oligopoly structure with three large national brand manufacturers operating on the upstream market (Nestlé, Danone, and Groupe Alma). Food retailers thus face a highly concentrated upstream market, a feature commonly shared with industries affected by the buyer alliances (e.g., soft drinks, beer, coffee). Second, given that over 80% of total bottled water sales take place in supermarket chains, the retail distribution channel stands as the primary outlet for the bottled water manufacturers. Third,

[22] Anecdotal evidence suggests that these negotiations are particularly fierce. For instance, retailers have been subject to recurrent legal actions in recent years (e.g., in 2015 the Paris Court of Appeal condemned Leclerc to reimburse more than €61 million to its suppliers for unfair practices in 2009 − 2010).

[23] The change in input price discrimination regulation and the long-run consequences of the economic crisis in 2008 have been advanced as explanatory factors for this wave of buyer consolidation (see Autorité de la concurrence, 2015). More specifically, it has been claimed that the aforementioned retailers have decided to join forces in negotiations to reduce their purchasing costs and maintain their competitiveness in the downstream market. In line with the empirical literature on mergers, my analysis abstracts away from these considerations and takes retailers' decisions to form buyer alliances as given.

bottled water ranks among the top-selling items for food retailers.[24]

3.2 Data

I use household-level scanner data including 550,059 purchases of bottled water in France collected by Kantar WorldPanel from March to December 2013 (pre-alliances) and March to December 2015 (post-alliances).[25] The data consist of a panel of households representative of the French population who record their grocery purchases for home consumption. Recorded information for each purchase of bottled water includes the purchase date, the quantity bought, the per-unit price of the bottled water (henceforth referred to as the retail price), and key attributes such as the brand name, the type of water (mineral, sparkling), and whether it is flavored or not. The data also provide details about the store in which each purchase was made such as its name, its size area, and its type (e.g., traditional food store, supermarket, hypermarket). I also have access to household characteristics such as the age of the household head or an income-reported interval for each household in the panel.

I focus on purchases of bottled water in stores with a floor area exceeding four hundred square meters and owned by either one of the seven largest retailers or a hard discounter. Among these bottled water purchases, I select the 11 most purchased national brands as well as all purchases of private labels. Each selected national brand is produced by one of the three manufacturers (Nestlé produces 5 brands, Danone produces 4 brands, and Groupe Alma produces 2 brands). Private labels are distinguished according to four types of bottled water (mineral or spring water and still or sparkling water) and their manufacturers are assumed to be vertically integrated with the retailers (i.e., there is no bargaining over their wholesale prices). Other purchases of bottled water in the sample are lumped together under the label "outside good" and include flavored water, national brands with a small purchase frequency, and bottled water purchased at small stores. Finally, I define a product as a brand-retailer combination, resulting in a total of 111 differentiated products.[26]

As in most revealed-preference data, I have no information on products other than those purchased by consumers during their shopping trips (that is, I do not observe the choice set available to each consumer). To address this issue, I first define the market as all purchases of bottled water for home consumption in France within a month. Then, I compute

[24]According to NielsenIQ, the top-selling item in supermarkets in 2022 is the brand of bottled water Cristaline (with Coca-Cola ranking second). Furthermore, 8 other national brands of bottled water (e.g., Volvic, Evian) feature among the top 50 best-selling items.

[25]I use the terms "household" and "consumer" interchangeably throughout this article.

[26]More precisely, I consider that a national brand sold by two retailers corresponds to two different products. Note that this definition aggregates different package sizes of bottled water.

Table 1: Descriptive statistics for manufacturers and retailers

	Market share		Retail price	
	pre-alliances	post-alliances	pre-alliances	post-alliances
Upstream level:				
Manufacturer 1	15.69 (1.07)	16.72 (0.48)	0.53 (0.02)	0.48 (0.01)
Manufacturer 2	10.86 (0.43)	10.57 (0.28)	0.46 (0.01)	0.43 (0.02)
Manufacturer 3	13.09 (0.76)	14.71 (0.95)	0.22 (0.00)	0.19 (0.01)
Private labels	23.40 (0.54)	21.31 (0.40)	0.26 (0.00)	0.26 (0.00)
Downstream level:				
Retailer 1	14.84 (0.36)	14.21 (0.51)	0.40 (0.01)	0.37 (0.01)
Retailer 2	1.78 (0.16)	1.86 (0.18)	0.43 (0.02)	0.39 (0.02)
Retailer 3	7.31 (0.43)	7.20 (0.48)	0.42 (0.01)	0.40 (0.01)
Retailer 4	4.94 (0.21)	5.99 (0.17)	0.40 (0.01)	0.37 (0.01)
Retailer 5	8.97 (0.76)	8.94 (0.40)	0.40 (0.01)	0.37 (0.01)
Retailer 6	4.62 (0.19)	4.75 (0.25)	0.42 (0.01)	0.39 (0.01)
Retailer 7	14.48 (0.67)	14.17 (0.53)	0.39 (0.01)	0.37 (0.01)
Retailer 8	6.10 (0.10)	6.19 (0.30)	0.26 (0.01)	0.26 (0.01)
Outside good	36.96 (1.52)	36.69 (0.78)	–	–

Notes: Market shares in percentage are calculated according to the number of household purchases. Average retail prices are in euros per liter. Standard deviations in parenthesis refer to variation across markets for the year 2013 (pre-alliances) and 2015 (post-alliances). Remark that I am not permitted to reveal names of manufacturers and retailers due to confidentiality regarding Kantar WorldPanel data.

a monthly average (deflated) retail price for each product.[27] Considering the most popular brands of bottled water sold by the largest retailers in France, I make the assumption that each consumer in the sample has made his purchasing decision among the 111 differentiated products (plus the outside good) sold at their respective monthly retail prices.

Table 1 reports descriptive statistics about the market shares and retail prices of products for each manufacturer and retailer before and after the formation of buyer alliances (see Appendix A for additional statistics). In the pre-alliances periods, the market shares of

[27] The retail price of product j in market t is constructed as follows: $p_{j,t} = \frac{\sum_i \mathbb{1}_{i,j,t} p_{i,j,t} q_{i,j,t}}{\sum_i \mathbb{1}_{i,j,t} q_{i,j,t}}$, where $\mathbb{1}_{i,j,t}$ is an indicator equals to 1 if consumer i has purchased product j in market t, $p_{i,j,t}$ is the retail price paid by the consumer, and $q_{i,j,t}$ is the quantity purchased (in liters).

national brand manufacturers range from 10.86% to 15.69%, and private labels account for 23.40% of total bottled water purchases. Private labels are on average sold at half the retail price of national brands of manufacturers 1 and 2. However, manufacturer 3's national brands are on average cheaper than private labels. This suggests that there is a strong heterogeneity between national brands in the retail price dimension.[28] In the post-alliances periods, the average retail price of private labels remains unchanged relative to the pre-alliances periods. In contrast, the average retail price of national brand manufacturers is substantially lower, reducing the price gap between national brands and private labels. The market shares of manufacturers 1 and 3 increase whereas sales of private labels decrease relative to the pre-alliances periods.

The table also shows substantial variation of market shares across retailers, ranging from 1.78% for the smallest to 14.84% for the largest retailer in the pre-alliances periods. As further shown in Table 8 of Appendix A, sales of national brands constitute the largest portion of retailers' market shares, except for retailer 8 whose total sales are composed at 87% of private labels. The retail prices of products sold by retailers 7 and 8, which have not formed any alliance in 2014, are on average lower than the retail prices charged by other retailers (see also Table 8 of Appendix A). This price gap, however, reduces in the post-alliances periods. In what follows, I consider a descriptive analysis of retail prices to explore the dynamic of this price variation.

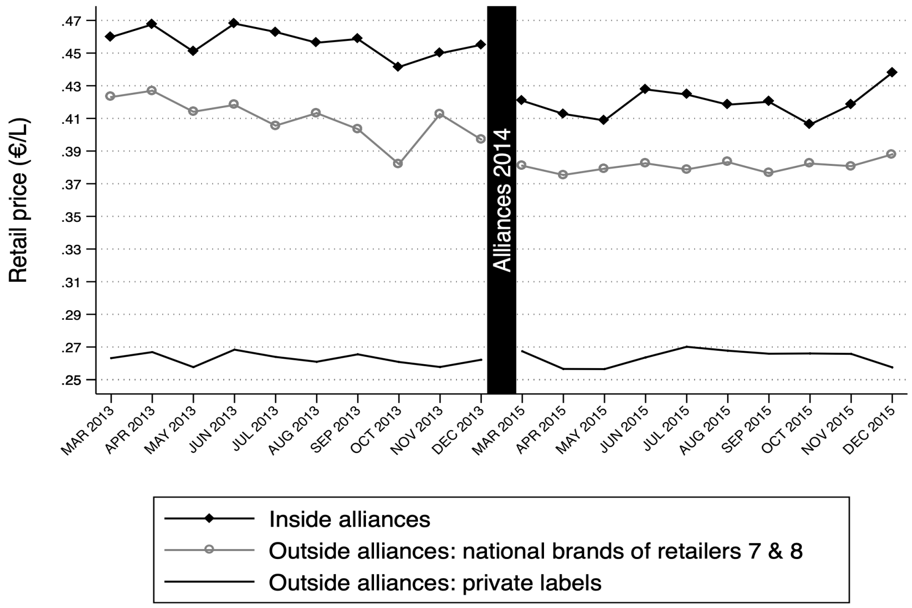
3.3 Descriptive retail price analysis

Figure 4 shows the time path of the average retail prices of three product groups: (i) national brands sold by retailers involved in a buyer alliance in 2014, (ii) national brands sold by retailers that have not formed any alliance in 2014 (retailers 7 and 8), and (iii) private labels which have all been excluded from the scope of the alliances. The figure shows that the average retail price of national brands sold by retailers involved in an alliance experienced a sharp decrease from about €0.46 to €0.42 after the formation of buyer alliances. Although a decline in the retail prices of national brands sold by retailers 7 and 8 is also observed in the year following the alliances, the average retail price of private labels remains stable over time.

To quantify this variation in retail prices, I conduct an event-study analysis. After collapsing the data at the product-market level, I estimate the following reduced-form pricing

[28]This heterogeneity may be partly explained by cost differences as manufacturer 3's leading brand is a spring water extracted from 32 underground sources throughout France.

Figure 4: Retail price trends



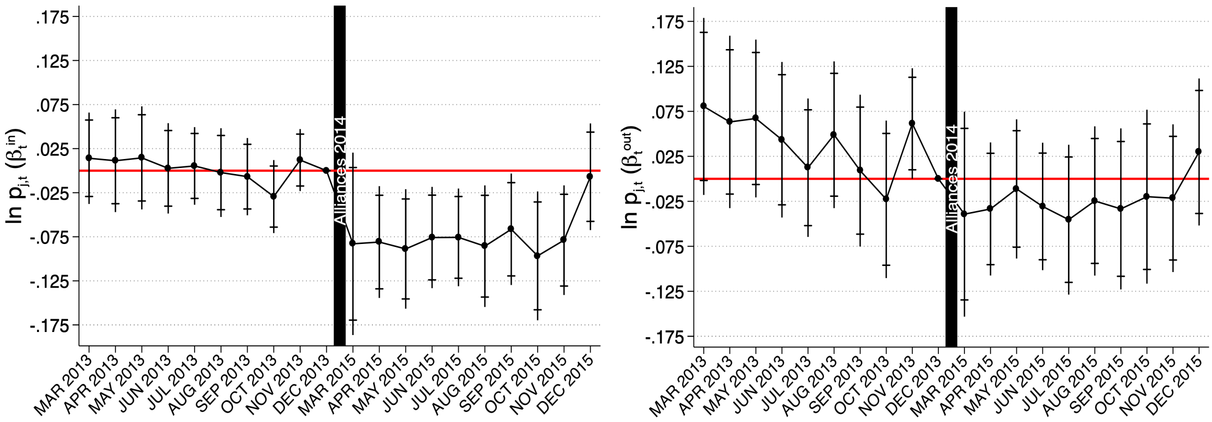
Notes: The black line with diamond markers is the average retail price trend of national brands sold by retailers that have formed a buyer alliance in 2014. The gray line with circle markers is the average retail price trend of national brands sold by retailers 7 and 8 which have not formed any alliance. The black line without markers is the average retail price trend of private labels sold by all retailers.

equation:

$$\begin{aligned} \ln p_{j,t} = \quad & \beta_j + \beta_t + \beta_t^{in} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \\ & + \beta_t^{out} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} + u_{j,t} \end{aligned} \tag{1}$$

where $\ln p_{j,t}$ is the natural log retail price of product j in market t, β_j and β_t are product and market fixed effects controlling for the cross-sectional and time variation in the retail prices of products, and $u_{j,t}$ is an error term capturing unobserved factors affecting retail prices. The indicator variables $\mathbb{1}\{\text{national brand}\}_{j,t}$, $\mathbb{1}\{\text{alliance}\}_{j,t}$, and $\mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t}$ are respectively equal to 1 if product j in market t is a national brand, if it is sold by a retailer involved in an alliance in 2014, and if it is sold by either retailer 7 or 8. The parameters of interest are the fixed effects β_t^{in} and β_t^{out}. They measure the evolution of the log retail price difference between national brands sold by a retailer involved in an alliance and private labels (β_t^{in}), and the evolution of the log retail price difference between national brands sold by retailer 7 or 8 and private labels (β_t^{out}). One can interpret these parameters as a measure of the effect of the alliances on retail prices under the two following assumptions. First, the retail prices of private labels have not been affected by the alliance formation. Second, private labels and national brands face comparable demand and supply conditions

Figure 5: Event study



(a) Inside alliances (b) Outside alliances: retailers 7 & 8

Notes: OLS estimates of the coefficients β_t^{in} and β_t^{out} from the regression model (1). The number of observations is 2,192 and the R^2 adjusted equals 0.99. Capped-bars and bars indicate respectively the 90% and 95% confidence intervals with standard errors clustered at the product level (111 clusters). The coefficients β_t^{in} and β_t^{out} for December 2013 are normalized to zero (base period). Percentage changes in retail prices can be obtained from the following transformation of the estimated parameters: $100(\exp(\beta)-1)$.

such that their retail prices would have followed the same trend absent buyer alliances.[29] I estimate equation (1) by ordinary least squares (OLS) with standard errors clustered at the product level.[30] Figures 5a and 5b display the estimation results (see Appendix B.2 for an alliance-by-alliance event-study analysis).

Figure 5a shows that the trend in retail prices of national brands sold by retailers involved in an alliance during the pre-alliances periods is stable and not statistically different from that of private labels. In the post-alliances periods, however, there is a clear trend break in which most parameters are significantly negative with values around −0.075. This suggests that, after the formation of buyer alliances, the average retail price of national brands sold by a member of an alliance has decreased by 7.23% relative to that of private labels. Figure 5b also indicates that the retail prices of national brands sold by retailers 7 and 8 do not trend differently from the retail prices of private labels in the pre-alliances periods. In the post-alliances periods, the estimates show a slight decrease in the retail prices of their national brands (approximately −2.50%). While these estimates suggest that retailers compete in strategic complements, the price response of retailers 7 and 8 to the formation of the alliances is not statistically significant.

This preliminary analysis provides descriptive evidence that the retail prices of national

[29] Although not perfect, private labels are often used as a control group in the retrospective merger literature (see, e.g., Ashenfelter and Hosken, 2010; Weinberg and Hosken, 2013).

[30] Clustering at more aggregated levels such as brand-retailer (89 clusters) or manufacturer-retailer-type of water (75 clusters) yields qualitatively similar results. In Appendix B.1, I also consider the "aggregation" approach suggested by Bertrand, Duflo and Mullainathan (2004) as an alternative solution for the serial correlation of retail prices.

brands sold by retailers involved in an alliance sharply decreased in the post-alliances periods. Several shortcomings are however worth mentioning. Although the event study reveals a strong correlation between the decrease in retail prices of national brands and the formation of buyer alliances, caution regarding its causal interpretation is warranted.[31] Furthermore, this analysis does not offer guidance on the mechanisms underlying changes in retail prices and their implications for the industry profit and its division between manufacturers and retailers, which are subjects of intense political and antitrust debates. In what follows, I develop a structural model of demand and supply to explicitly address these issues.

4 Empirical framework

This section introduces the structural model of demand and supply that I take to data. I first describe the demand-side which models consumer choices for bottles of water in supermarket chains. Then, I turn to the supply-side which models the bottled water industry as a bilateral oligopoly.

4.1 Consumer demand for bottled water

4.1.1 Demand model

I consider a demand system that derives from a standard discrete choice model of consumer behavior. More specifically, I use a random coefficient logit model that accommodates rich substitution and curvature patterns through observed and unobserved heterogeneity in consumer preferences.

Suppose that each consumer $i = 1, \dots, I_t$ in the sample chooses among $J_t + 1$ alternatives indexed from $j \in \{0, \dots, J_t\} \equiv \mathscr{J}_t$ on each shopping trip in market t. Alternative $j = 0$ is referred to as the composite "outside good", while other alternatives correspond to J_t products called "inside goods".[32] Each inside good j is associated with a brand $b = 1, \dots, B$ — where $b(j)$ labels the brand of good j — sold by a retailer $r = 1, \dots, R$ — where $r(j)$ denotes the retailer distributing good j.

The indirect utility function of consumer i from purchasing inside good j in market t is

[31]Ideally, I would use as a control group the same national brands sold by retailers in adjacent markets where the buyer alliances did not occur. However, the national scope of the alliances precludes the availability of such products in my data.

[32]Terms "good" and "product" are used interchangeably and refer to alternatives in the choice set $\mathscr{J}_t$.

specified as follows:

$$U_{i,j,t} = \mathbf{x}_j^\top \boldsymbol{\beta}_i - \alpha_i p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t} + \epsilon_{i,j,t} \tag{2}$$

where $\mathbf{x}_j$ is a L-dimensional vector of product characteristics that includes indicator variables for sparkling water, mineral water, and private label (the superscript "$\top$" denotes the transpose operator). It also includes a constant term that captures the utility generated by time-invariant characteristics common to all inside goods. The terms δ_t are market fixed effects controlling for changes over time in the valuation of characteristics common to all inside goods, $\delta_{b(j)}$ and $\delta_{r(j)}$ are respectively national brand and retailer fixed effects which capture the mean valuation of time-invariant characteristics specific to each brand and each retailer, $\xi_{j,t}$ is a structural error which embeds the mean utility generated by product characteristics unobserved to the econometrician, and $\epsilon_{i,j,t}$ is a stochastic term representing unobserved consumer-specific preferences. Preferences for the time-invariant characteristics common to all inside goods, mineral water, sparkling water, and private labels ($\boldsymbol{\beta}_i$) as well as the price sensitivity (α_i) are allowed to vary across consumers. More specifically, I assume that these preference parameters are distributed as follows:

$$\alpha_i = \alpha + \sigma^o_{g(i),p} d^y_{g(i)} + \sigma^u_p \nu_{i,p} \quad \text{and} \quad \boldsymbol{\beta}_{i,l} = \begin{cases} \beta_l + \sigma^o_{g(i),l} d^a_{g(i)} + \sigma^u_l \nu_{i,l} & \text{if } l = \text{mineral} \\ \beta_l + \sigma^u_l \nu_{i,l} & \text{otherwise} \end{cases}$$

where α and β_l capture respectively the mean retail price sensitivity and the mean valuation of consumers for the lth product characteristic. The retail price sensitivity and the valuation for mineral water are allowed to vary across consumers according to their observed income group ($d^y_{g(i)}$) and age group ($d^a_{g(i)}$) respectively.[33] Hence, the parameters $\sigma^o_{g(i),p}$ and $\sigma^o_{g(i),\text{mineral}}$ measure the degree of observed heterogeneity in the retail price sensitivity and the valuation of consumers for mineral water. Consumer preferences are also allowed to vary according to some unobserved consumer characteristics represented by $\nu_{i,p}$ and $\nu_{i,l}$, which are assumed to be independent and identically distributed standard normal. Hence, the σ^u parameters measure the degree of unobserved heterogeneity in the retail price sensitivity and the valuation of consumers for product characteristics. Such parameters governing the degree of heterogeneity in consumer preferences are likely to play an important role in the analysis of buyer alliances. Indeed, they allow for rich substitution and curvature patterns, which have direct implications for both the status quo and the nondiscrimination

[33] I consider four monthly income groups (below € 900, between € 900 and € 1,899, between € 1,900 and € 4,449, and above € 4,449) and three age groups (below 40, between 40 and 60, and above 60).

effects of buyer alliances described in Section 2.

The indirect utility that consumer i receives from choosing the outside good in market t is normalized as follows: $U_{i,0,t} = \epsilon_{i,0,t}$. Assuming that each consumer in the sample is a utility maximizer (i.e., he chooses one unit of the good that gives him the highest utility) and that $\epsilon_{i,j,t}$ is independently and identically distributed from the standard Gumbel distribution (also known as type I extreme value distribution), the probability that consumer i selects product $j \in \mathcal{J}_t \backslash \{0\}$ in market t is given by:

$$\mathcal{s}_{i,j,t} = \int_{\mathbb{R}^{L+1}} \frac{\exp(\mathbf{x}_j^\top \boldsymbol{\beta}_i - \alpha_i p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t})}{\sum_{k \in \mathcal{J}_t} \exp(\mathbf{x}_k^\top \boldsymbol{\beta}_i - \alpha_i p_{k,t} + \delta_{b(k)} + \delta_{r(k)} + \delta_t + \xi_{k,t})} f(\boldsymbol{\nu}_i) \mathrm{d}\boldsymbol{\nu}_i \tag{3}$$

where $\boldsymbol{\nu}_i \equiv (\nu_{i,p}, \nu_{i,\text{constant}}, \nu_{i,\text{mineral}}, \nu_{i,\text{sparkling}}, \nu_{i,\text{private label}})^\top$ and $f(\cdot)$ denotes the joint normal distribution function. Following Revelt and Train (1998), the probability of observing the sequence of consumer i's choices in the panel data is:

$$\mathcal{S}_i = \int_{\mathbb{R}^{L+1}} \prod_{t=1}^{T} \prod_{j \in \mathcal{J}_t} \left(\frac{\exp(\mathbf{x}_j^\top \boldsymbol{\beta}_i - \alpha_i p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t})}{\sum_{k \in \mathcal{J}_t} \exp(\mathbf{x}_k^\top \boldsymbol{\beta}_i - \alpha_i p_{k,t} + \delta_{b(k)} + \delta_{r(k)} + \delta_t + \xi_{k,t})} \right)^{\mathbb{1}_{i,j,t}} f(\boldsymbol{\nu}_i) \mathrm{d}\boldsymbol{\nu}_i \tag{4}$$

where $\mathbb{1}_{i,j,t}$ is an indicator variable equals to 1 if consumer i purchases product j in market t and 0 otherwise. In what follows, I consider the identification and estimation of the vector of demand parameters $\boldsymbol{\theta}^d \equiv (\boldsymbol{\beta}, \alpha, \boldsymbol{\delta}_{\text{brand}}, \boldsymbol{\delta}_{\text{retail}}, \boldsymbol{\delta}_{\text{market}}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u)^\top$.[34]

4.1.2 Identification and estimation of consumer demand

Identification assumptions. As stressed in Berry (1994), identification of $\boldsymbol{\theta}^d$ can be jeopardized by the classical endogeneity problem of the retail price variable. For instance, whenever retailers observe the realization of demand shocks for unobserved product characteristics before setting retail prices, $p_{j,t}$ is likely to be correlated with $\xi_{j,t}$. To address this issue and obtain consistent estimates of consumer preferences, I follow Berry, Levinsohn and Pakes (2004). Let $\delta_{j,t} \equiv \mathbf{x}_j^\top \boldsymbol{\beta} - \alpha p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t}$ subsumes the mean utility level of product $j \in \mathcal{J}_t \backslash \{0\}$. Combined with the micro data, this approach allows identifying the vector of mean utility levels $\boldsymbol{\delta} \equiv (\delta_{1,1}, \ldots, \delta_{J_T,T})^\top$ as well as $\boldsymbol{\sigma}^o$ and $\boldsymbol{\sigma}^u$ without restriction on the distribution of the structural error $\xi_{j,t}$. Note that among this first set of parameters,

[34]The vector $\boldsymbol{\beta}$ includes the constant term and the parameters that capture the mean valuation of consumers for mineral water, sparkling water, and private labels. The vectors $\boldsymbol{\delta}_{\text{brand}}$, $\boldsymbol{\delta}_{\text{retail}}$, and $\boldsymbol{\delta}_{\text{market}}$ include respectively all national brand, retailer, and market fixed effects. The vectors $\boldsymbol{\sigma}^o$ and $\boldsymbol{\sigma}^u$ include respectively all parameters that govern the observed and unobserved heterogeneity in consumer preferences.

identification of the degree of unobserved heterogeneity in consumer preferences ($\boldsymbol{\sigma}^u$) is the most subtle. I leverage the panel structure of the micro data which, through the correlation in the characteristics of products purchased by each consumer over time, provides a useful source of information on $\boldsymbol{\sigma}^u$.[35]

Identification of the mean taste parameters $\boldsymbol{\beta}$, α, $\boldsymbol{\delta}_{\text{brand}}$, $\boldsymbol{\delta}_{\text{retail}}$, and $\boldsymbol{\delta}_{\text{market}}$ requires assumptions on the distribution of $\xi_{j,t}$. I rely on an instrumental variable approach consisting of finding at least one variable that affects $p_{j,t}$ by shifting supply (costs or markups of retailers) but not preferences for unobserved characteristics of bottled water. I consider two instrumental variables. Under the assumption that the observed product characteristics are uncorrelated with $\xi_{j,t}$ (Berry, Levinsohn and Pakes, 1995), I use the number of products other than j sold by the same retailer in each market ($\sum_{k \in \mathcal{J}_{r(j),t} \setminus \{j\}} 1$).[36] The second instrument is based on insights developed in Berry and Haile (2014) and applied by Miller and Weinberg (2017) to the analysis of horizontal mergers in oligopoly. By interpreting the formation of buyer alliances as an exogenous shifter of the bargaining environment in the vertical chain, I use the indicator variable $\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$ as instrument for $p_{j,t}$.[37] Given the presence of brand, retailer, and market fixed effects in (2), exogeneity of this instrument relies on the assumption that changes in the structural error of national brands sold by retailers involved in an alliance, before versus after the alliances, are not systematically different from changes in the structural error of national brands sold by retailers 7 and 8 and private labels.[38]

Estimation procedure. I follow the spirit of the two-step estimation procedure of Berry, Levinsohn and Pakes (2004). Using consumer-level panel data, I first estimate the vectors of parameters $\boldsymbol{\delta}$, $\boldsymbol{\sigma}^o$, and $\boldsymbol{\sigma}^u$ by simulated maximum likelihood (SML). Formally, using Monte Carlo integration to simulate the integral in (4), the SML estimator is defined as follows:

$$(\boldsymbol{\delta}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u) \equiv \underset{\boldsymbol{\delta}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u}{\text{argmax}} \sum_{i=1}^{I} \ln\left(\tilde{\mathcal{S}}_i(\boldsymbol{\delta}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u)\right) \tag{5}$$

[35] For instance, observing that some consumers always purchase cheap products whereas others always purchase expensive products helps reveal the degree of unobserved heterogeneity in the retail price sensitivity.

[36] The main motive for this instrument is that retail prices set by a retailer depend on the number of products it distributes (e.g., a retailer with a large set of products is more likely to set high retail prices).

[37] This indicator can be interpreted as a cost shifter for retailers given that the primary motive for the alliances is to reduce wholesale prices paid by retailers to manufacturers. The descriptive retail price analysis in Section 3.3 provides direct evidence of the relevance of this instrument.

[38] This assumption would be violated if, for instance, *every* retailer involved in an alliance reorganizes its retail services (e.g., change in shelf display), thereby affecting consumer preferences for unobserved attributes in the post-alliances periods. Given that the focus of the alliances is on the negotiation of wholesale prices, I can reasonably assume that such a coordinated change in retail services is unlikely to occur.

where I denotes the total number of consumers in the sample. The term $\tilde{\mathcal{S}}_i(\boldsymbol{\delta}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u)$ in (5) is the simulated counterpart of (4), that is:

$$\tilde{\mathcal{S}}_i(\boldsymbol{\delta}, \boldsymbol{\sigma}^o, \boldsymbol{\sigma}^u) = \frac{1}{R}\sum_{r=1}^{R}\prod_{t=1}^{T}\prod_{j\in\mathcal{J}_t}\left(\frac{\exp(\delta_{j,t} + \upsilon_{i,r,j,t})}{\sum_{k\in\mathcal{J}_t}\exp(\delta_{k,t} + \upsilon_{i,r,k,t})}\right)^{\mathbb{1}_{i,j,t}}$$

where $\upsilon_{i,r,j,t} \equiv \sigma^o_{g(i),p} d^y_{g(i)} p_{j,t} + \sigma^u_p p_{j,t} \nu_{i,r,p} + \sigma^o_{g(i),\text{mineral}} d^a_{g(i)} x_{j,\text{mineral}} + \sum_l \sigma^u_l x_{j,l} \nu_{i,r,l}$ and R is the total number of draws from the joint normal distribution $f(\cdot)$. In practice, I use $R = 200$ shuffled Halton draws for each consumer i in the sample.[39] I refer to Section S3 of the Supplemental Material for further details about the estimation procedure.

For the second step, I use the estimated $\hat{\boldsymbol{\delta}}$ and consider the following linear regression model: $\hat{\boldsymbol{\delta}} = \mathbf{x}_j^\top \boldsymbol{\beta} - \alpha p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t}$. The mean taste parameters $\boldsymbol{\beta}$, α, $\boldsymbol{\delta}_{\text{brand}}$, $\boldsymbol{\delta}_{\text{retail}}$, and $\boldsymbol{\delta}_{\text{market}}$ are estimated using a two-stage least square estimator (TSLS), where instrumental variables are the number of other products sold by the same retailer in each market and the indicator variable for the formation of buyer alliances.

4.2 Downstream competition and manufacturer-retailer bargaining

I model the French bottled water industry as a bilateral oligopoly. In each market t, F multi-product manufacturers deal with R multi-product retailers to supply their products to consumers. Let $\mathcal{J}_{f,t}$ be the set of products owned by manufacturer f and $\mathcal{J}_{r,t}$ be the set of products distributed by retailer r in market t such that $\bigcup_{f=1}^{F} \mathcal{J}_{f,t} = \bigcup_{r=1}^{R} \mathcal{J}_{r,t} = \mathcal{J}_t \backslash \{0\}$. Define respectively the (per-market) profit function of manufacturer f and retailer r as follows:

$$\pi_{f,t} \equiv \sum_{j\in\mathcal{J}_{f,t}} \left(w_{j,t} - \mu_{b(j),t}\right) M_t s_{j,t}(\mathbf{p}_t, \boldsymbol{\theta}^d) \tag{6a}$$

$$\pi_{r,t} \equiv \sum_{j\in\mathcal{J}_{r,t}} \left(p_{j,t} - w_{j,t} - c_{j,t}\right) M_t s_{j,t}(\mathbf{p}_t, \boldsymbol{\theta}^d) \tag{6b}$$

where $w_{j,t}$ is the wholesale price of product j in market t, $\mu_{b(j),t}$ and $c_{j,t}$ are respectively the constant marginal cost of production for brand *b(j)* and the constant marginal cost of distribution for product j in market t,[40] M_t denotes the total quantity purchased in market t

[39] The shuffled Halton draws typically outperform the standard Halton draws for simulation of multi-dimensional integrals (e.g., Hess, Polak and Daly, 2003).

[40] I make the simplifying assumption that marginal costs of production only vary across brands of bottled water. This allows me to identify the price-cost margins of manufacturers in the post-alliances periods where manufacturers cannot price-discriminate between retailers of a buyer alliance (details are provided in Section S4.2.2 of the Supplemental Material). See Villas-Boas (2009) for a similar assumption. Although this

("market size"), and $s_{j,t}$ is the predicted market share of product j in market t written as a function of retail prices — denoted by the J_t-dimensional vector $\mathbf{p}_t$ — and demand parameters θ^d.[41]

Timing, information, and equilibrium concept. I consider a two-stage game in which firms interact as follows. In the first stage, manufacturers and retailers engage in simultaneous and secret bilateral negotiations over linear wholesale prices of products. In the second stage, retailers compete in retail prices on the downstream market.

As outlined in Section 2 and following recent empirical work on bargaining with externalities (e.g., Crawford and Yurukoglu, 2012; Grennan, 2013; Gowrisankaran, Nevo and Town, 2015; Ho and Lee, 2017), I employ the "Nash-in-Nash" bargaining solution to determine the division of surplus between manufacturers and retailers. Moreover, due to contract secrecy, I consider that retailers are not able to observe trading terms negotiated by their rivals before setting retail prices on the downstream market.[42] Assuming complete information about the cost of production and distribution for each product, I solve this two-stage game proceeding backwards.

To gain insights into the equilibrium notion of the above two-stage game, I refer to Rey and Vergé (2020) who provide a microfoundation for the "Nash-in-Nash" bargaining solution in a model of vertical relations with linear wholesale contracts. More specifically, they show that the above game replicates the sequential equilibrium (Kreps and Wilson, 1982) of a two-stage noncooperative game that includes a model of delegated negotiations in the first stage, followed by a model of retail price competition similar to my second stage.[43]

Remark on the sequential timing. As in Bonnet, Bouamra-Mechemache and Molina (forthcoming), I consider a two-stage game in which wholesale prices are determined *before* retail prices. It is worth mentioning that such a sequential timing differs from most empirical frameworks of bilateral oligopoly which, mainly for tractability reasons, consider instead

restriction is maintained in my notations, it has no implication for the identification of the price-cost margins of manufacturers in the pre-alliances periods.

[41] The market share of product $j \in \mathcal{J}_t$ is given by $s_{j,t} = \int_{\mathbb{R}^{L+1}} \frac{\exp(\mathbf{x}_j^\top \boldsymbol{\beta}_i - \alpha_i p_{j,t} + \delta_{b(j)} + \delta_{r(j)} + \delta_t + \xi_{j,t})}{\sum_{k \in \mathcal{J}_t} \exp(\mathbf{x}_k^\top \boldsymbol{\beta}_i - \alpha_i p_{k,t} + \delta_{b(k)} + \delta_{r(k)} + \delta_t + \xi_{k,t})} f(\boldsymbol{\nu}_i) d\boldsymbol{\nu}_i$ (see, e.g., Nevo, 2001).

[42] This information structure is often referred to as "interim unobservability" in the vertical contracting literature (see, e.g., Rey and Vergé, 2004; Gaudin, 2019).

[43] Collard-Wexler, Gowrisankaran and Lee (2019) provide an alternative noncooperative foundation for the "Nash-in-Nash" bargaining solution that relaxes delegation under the assumption that wholesale tariffs do not affect the total gains from trade generated by bilateral agreements (e.g., lump-sum payments).

that wholesale and retail prices are set simultaneously (e.g., Draganska, Klapper and Villas-Boas, 2010; Ho and Lee, 2017; Crawford et al., 2018). The sequential timing adopted here is in line with recent evidence that retail prices in the supermarket industry immediately adjust to wholesale price changes (see, e.g., Nakamura and Zerom, 2010; Goldberg and Hellerstein, 2013).

Remark on the contractual form. I consider negotiations over simple linear wholesale prices, which gives rise to the presence of double marginalization. To motivate this modeling assumption over more sophisticated contractual forms, I rely on Section 3.3 which provides descriptive evidence that retail prices of national brands substantially declined following the year of the alliances.[44] Under nonlinear tariffs, this variation can only be rationalized through cost savings caused by the alliances.[45] Under linear tariffs, however, a change in the distribution of bargaining power that mitigates double marginalization can also explain this retail price drop. I argue that the cost savings argument is implausible for at least two reasons. First, as described in Section 3.1, the alliances cover only the wholesale price negotiations of national brands and do not aim to restructure the distribution system of retailers. Second, empirical studies on mergers have documented that cost savings are long-run consequences and take roughly two years after a merger to materialize (e.g., Focarelli and Panetta, 2003; Ashenfelter, Hosken and Weinberg, 2015), which is well beyond the time period covered by my data. Hence, under the assumption that there are no unobserved demand or cost shocks unrelated to the formation of buyer alliances that affect only national brands purchased by an alliance, linear wholesale prices are arguably more appropriate to explain the variation observed in the data than nonlinear tariffs that would give rise to cost-based marginal wholesale prices.[46]

[44] Luco and Marshall (2020) rely on a similar approach to infer the contractual form used in vertical relations. Leveraging the variation created by vertical mergers, they find evidence in favor of a double marginalization in the soft drink industry, which typically arises under linear wholesale prices.

[45] This reasoning is based on Rey and Vergé (2020) who show in a framework similar to mine that equilibrium nonlinear tariffs are necessarily cost-based (i.e., the per-unit prices reflect marginal costs of production). Retail prices thus simply replicate the outcome of a multi-product oligopoly and are never affected by the distribution of bargaining power in the vertical chain. In addition, Caprice and Rey (2015) show that the formation of a buyer alliance under two-part tariffs is either neutral or increases retail prices by solving the manufacturer's opportunism problem (e.g., McAfee and Schwartz, 1994). Therefore, absent cost savings, theoretical predictions under nonlinear tariffs are in stark contrast with the variation observed in my data.

[46] This finding contrasts with Bonnet and Dubois (2010) who provide empirical evidence in favor of two-part tariff contracts with resale price maintenance under the assumption that manufacturers make public take-it-or-leave-it offers to retailers. It is worth noting, however, that their analysis of the French bottled water industry focuses on the years 1998 to 2000 where input price discrimination was banned and the Galland Act prevented retailers from resale below the per-unit invoice price set by manufacturers (but "off-invoice" rebates were allowed). Because these rules generated inflationary effects, the Galland Act has been repealed and input price discrimination allowed to reinvigorate retail competition, implying that the regulatory environment during the

4.2.1 Stage 2: Downstream competition

I assume that retail prices are determined in a pure-strategy Nash equilibrium where each retailer holds consistent beliefs about the wholesale contracts formed by manufacturers and its downstream rivals.[47] Using (6b), the maximization problem of retailer r in market t is given by $\max_{\{p_{j,t}\}_{j\in\mathscr{J}_{r,t}}} \pi_{r,t}$, implying that the pricing behavior of retailers is characterized by the following system of first-order conditions:

$$s_{k,t}(\mathbf{p}_t;\theta^d) + \sum_{j\in\mathscr{J}_{r(k),t}} \left(p_{j,t} - w_{j,t} - c_{j,t}\right) \frac{\partial s_{j,t}}{\partial p_{k,t}}(\mathbf{p}_t;\theta^d) = 0 \quad \forall k \in \mathscr{J}_t \backslash \{0\} \tag{7}$$

Based on (7), Berry and Haile (2014) show that identification of the price-cost margins and the marginal costs of retailers directly follows from identification of the demand model described in Section 4.1.[48] Importantly, there are no restrictions imposed on the structure of retailers' constant marginal costs. By inverting (7), I show in Section S4.1 of the Supplemental Material that one can recover the J_t-dimensional vector of price-cost margins of retailers denoted by $\boldsymbol{\gamma}_t$. The jth element of $\boldsymbol{\gamma}_t$ is given by $p_{j,t} - w_{j,t} - c_{j,t} = \gamma_{j,t}(\mathbf{S}_{\mathbf{p}_t}, s_t)$, where $\mathbf{S}_{\mathbf{p}_t}$ is the $J_t \times J_t$ matrix of first partial derivatives $\frac{\partial s_{k,t}}{\partial p_{j,t}}$ and s_t is the J_t-dimensional vector of market shares. Based on this result, I recover the marginal cost of retailers for each product $j \in \mathscr{J}_t \backslash \{0\}$ as follows: $w_{j,t} + c_{j,t} = p_{j,t} - \gamma_{j,t}$. As shown in Section 4.2.4, it turns out that retailers' marginal costs are key ingredients to the identification of firms' bargaining power in the vertical chain.

4.2.2 Stage 1: Manufacturer-retailer bargaining in the pre-alliances periods

In the pre-alliances periods, manufacturers and retailers engage in bilateral negotiations to determine the wholesale prices of products. The resolution of this stage can be described as follows.

Bargaining between manufacturer f and retailer r over $w_{j,t}$. Consider the bilateral negotiation between manufacturer f and retailer r over the wholesale price $w_{j,t}$, where $j \in$

time period covered by my data radically differs from their analysis.

[47] Following Rey and Vergé (2020), belief consistency implies here that a retailer anticipates in any circumstance that its rivals have negotiated the equilibrium wholesale prices (i.e., retailers believe that any out-of-equilibrium event is the result of a tremble).

[48] Following Berry and Haile (2014), two conditions ensure invertibility of (7) and, in turn, identification of retailers' marginal costs: (i) the differentiability of market shares with respect to retail prices, and (ii) that some products are strict substitutes in retail prices while others are weak substitutes from one another. The discrete choice model described in Section 4.1 belongs to the class of demand models satisfying these conditions.

$\mathscr{J}_{f,t} \cap \mathscr{J}_{r,t}$. Following previous assumptions, terms of trade are determined by the Nash bargaining solution taking as given the outcomes of other negotiations and anticipating the impact on retailer r's pricing behavior in stage 2. Hence, the equilibrium wholesale price $w^*_{j,t}$ solves the following Nash bargaining problem:

$$w^*_{j,t} \equiv \underset{w_{j,t}}{\text{argmax}} \left(\pi_{f,t} - d^{-j}_{f,t}\right)^{\lambda^{pre}_{f,r}} \left(\pi_{r,t} - d^{-j}_{r,t}\right)^{1-\lambda^{pre}_{f,r}} \tag{8}$$

where $\lambda^{pre}_{f,r} \in [0,1]$ corresponds to the bargaining weight of manufacturer f when negotiating with retailer r in the pre-alliances periods, and $\pi_{f,t}$ and $\pi_{r,t}$ are the profit of firms defined in (6a) and (6b). The terms $d^{-j}_{f,t}$ and $d^{-j}_{r,t}$ denote respectively the status quo payoffs of manufacturer f and retailer r in the event of bilateral disagreement:

$$\begin{aligned} d^{-j}_{f,t} &= \sum_{k \in \mathscr{J}_{f,t} \setminus \{j\}} \left(w^*_{k,t} - \mu_{b(k),t}\right) M_t \tilde{\mathfrak{z}}^{-j}_{k,t}(\tilde{\mathbf{p}}^{-j}_t; \theta^d) \\ d^{-j}_{r,t} &= \sum_{k \in \mathscr{J}_{r,t} \setminus \{j\}} \left(\tilde{p}^{-j}_{k,t} - w^*_{k,t} - c_{k,t}\right) M_t \tilde{\mathfrak{z}}^{-j}_{k,t}(\tilde{\mathbf{p}}^{-j}_t; \theta^d) \end{aligned}$$

where $w^*_{k,t}$ is the (anticipated) equilibrium wholesale price of product $k \neq j$, $\tilde{\mathbf{p}}^{-j}_t$ is the J_t-dimensional vector of out-of-equilibrium retail prices when product j is no longer offered on market t, and $\tilde{\mathfrak{z}}^{-j}_{k,t}$ is the market share of each product k remaining on the market (see Section S4.2.1 of the Supplemental Material for further details). Hence, this specification implies that each manufacturer-retailer pair negotiates the wholesale prices of products separately.[49] From (8), I can derive the set of first-order conditions characterizing the surplus division in each bilateral negotiation in market t:

$$\underbrace{\left(1 - \lambda^{pre}_{f(j),r(j)}\right)\left(\pi_{f(j),t} - d^{-j}_{f(j),t}\right)\frac{\partial \pi_{r(j),t}}{\partial w_{j,t}}}_{\text{retailer } r(j)\text{'s bargaining power}} + \underbrace{\lambda^{pre}_{f(j),r(j)}\left(\pi_{r(j),t} - d^{-j}_{r(j),t}\right)\frac{\partial \pi_{f(j),t}}{\partial w_{j,t}}}_{\text{manufacturer } f(j)\text{'s bargaining power}} = 0 \;\; \forall j \in \mathscr{J}_t \setminus \{0\} \tag{9}$$

where $f(j)$ indexes the manufacturer of product j. As discussed in Section 2, (9) shows that the bargaining power of a firm involved in a bilateral negotiation depends on three different factors. The first factor corresponds to the firm's bargaining ability vis-à-vis its trading partner, which is captured by the bargaining weight $\lambda^{pre}_{f(j),r(j)}$. The second is the (incremental) gains from trade obtained by the firm's trading partner, which derive from the difference between the trading partner's profits when all agreements are formed and its status quo pay-

[49] I have also considered an alternative specification with joint negotiations (see Section S8 of the Supplemental Material).

offs. The third factor is the cost incurred from making price concessions, which is captured by the derivative of the firm's profits with respect to the wholesale price.

By inverting the system (9) in a similar manner as for the inversion of (7) to recover the price-cost margins of retailers, I can obtain a closed-form expression for the J_t-dimensional vector of manufacturers' price-cost margins in the pre-alliances periods, denoted by $\mathbf{\Gamma}_t^{pre}$ (see Section S4.2.1 of the Supplemental Material). The jth element of $\mathbf{\Gamma}_t^{pre}$ is given by:

$$w_{j,t} - \mu_{b(j),t} = \Gamma_{j,t}^{pre}(s_t, \mathbf{S}_{\mathbf{p}_t}, \mathbf{S}_{\mathbf{p}_t\mathbf{p}_t}, \tilde{\mathbf{S}}_{\Delta t}^{pre}; \boldsymbol{\lambda}^{pre}) \tag{10}$$

where $\boldsymbol{\lambda}^{pre}$ is a J_t-dimensional vector of manufacturers' bargaining weight with $\boldsymbol{\lambda}^{pre}[j,1] = \lambda_{f(j),r(j)}^{pre}$, $\mathbf{S}_{\mathbf{p}_t\mathbf{p}_t}$ is an array of J_t matrices of second partial derivatives $\frac{\partial^2 s_{j,t}}{\partial p_{k,t}\partial p_{l,t}}$ denoted by $\mathbf{S}_{\mathbf{p}_{k,t}\mathbf{p}_t}$ (each matrix $\mathbf{S}_{\mathbf{p}_{k,t}\mathbf{p}_t}$ being of $J_t \times J_t$ dimension), and $\tilde{\mathbf{S}}_{\Delta t}^{pre}$ is a $J_t \times J_t$ matrix of difference in market shares $s_{k,t}(\mathbf{p}_t; \theta^d) - \tilde{s}_{k,t}^{-j}(\tilde{\mathbf{p}}_t^{-j}; \theta^d)$.

4.2.3 Stage 1: Manufacturer-retailer bargaining in the post-alliances periods

In the post-alliances periods, each manufacturer engages in bilateral negotiations with alliances of retailers to determine the wholesale prices of its products. Before considering the resolution of this bargaining stage, I introduce the following notations. Let *a(j)* denote the alliance of retailers purchasing product j,[50] $\mathcal{J}_{a,t}$ is the set of products purchased by the alliance a, and $\mathcal{J}_{b,t}$ is the set of products sold under the brand b. In line with Section 2, I consider that a buyer alliance negotiates on behalf of its members and aims at maximizing their joint profits. Furthermore, retailers involved in the alliance a obtain nondiscriminatory trading terms when purchasing brand b: e.g., $w_{j,t} = w_{k,t} = w_{a,b,t}$ $\forall j,k \in \mathcal{J}_{a,t} \cap \mathcal{J}_{b,t}$.[51]

Bargaining between manufacturer f and the buyer alliance a over $w_{j,t}$. Consider the bilateral negotiation between manufacturer f (offering brand b) and the alliance a over the wholesale price of product $j \in \mathcal{J}_{a,t} \cap \mathcal{J}_{b,t}$ ($w_{j,t} = w_{a,b,t}$). Taking as given the outcomes of other negotiations and anticipating the impact on the pricing behavior of *every* retailer involved in the alliance a in stage 2, the equilibrium wholesale price $w_{a,b,t}^*$ solves the following Nash bargaining problem:

$$w_{a,b,t}^* \equiv \underset{w_{a,b,t}}{\text{argmax}} \left(\pi_{f,t} - d_{f,t}^{-a,b}\right)^{\lambda_{f,a}^{post}} \left(\pi_{a,t} - d_{a,t}^{-a,b}\right)^{1-\lambda_{f,a}^{post}} \tag{11}$$

[50]If retailer *r(j)* has not formed any alliance with another retailer, I consider that *a(j)* = *r(j)*.

[51]To clarify notations, note that $w_{a(k),b(k),t} \neq w_{a,b,t}$ if product $k \notin \mathcal{J}_{a,t} \cap \mathcal{J}_{b,t}$. Indeed, $w_{a(k),b(k),t}$ denotes the wholesale price paid by the alliance *a(k)* (which may differ from the alliance a) for a product sold under the brand $b(k)$ (which may differ from brand b).

where $\lambda_{f,a}^{post} \in [0,1]$ denotes the bargaining weight of manufacturer f when negotiating with the alliance a in the post-alliances periods, and $\pi_{f,t}$ and $\pi_{a,t}$ are respectively the profit of manufacturer f and the alliance a when all agreements are formed:

$$\pi_{f,t} = \sum_{h \in \mathscr{J}_{a,t} \cap \mathscr{J}_{b,t}} \left(w_{a,b,t} - \mu_{b,t}\right) M_t s_{h,t}(\mathbf{p}_t; \theta^d) + \sum_{k \in \mathscr{J}_{f,t} \setminus \{\mathscr{J}_{a,t} \cap \mathscr{J}_{b,t}\}} \left(w^*_{a(k),b(k),t} - \mu_{b(k),t}\right) M_t s_{k,t}(\mathbf{p}_t; \theta^d)$$

$$\pi_{a,t} = \sum_{h \in \mathscr{J}_{a,t} \cap \mathscr{J}_{b,t}} \left(p_{h,t} - w_{a,b,t} - c_{h,t}\right) M_t s_{h,t}(\mathbf{p}_t; \theta^d) + \sum_{k \in \mathscr{J}_{a,t} \setminus \mathscr{J}_{b,t}} \left(p_{k,t} - w^*_{a,b(k),t} - c_{k,t}\right) M_t s_{k,t}(\mathbf{p}_t; \theta^d)$$

where $w^*_{a(k),b(k),t}$ corresponds to the (anticipated) equilibrium wholesale price for brand $b(k)$ paid by the alliance $a(k)$. The terms $d_{f,t}^{-a,b}$ and $d_{a,t}^{-a,b}$ represent respectively the status quo payoffs of manufacturer f and the alliance a in the event of bilateral disagreement:

$$d_{f,t}^{-a,b} = \sum_{k \in \mathscr{J}_{f,t} \setminus \{\mathscr{J}_{a,t} \cap \mathscr{J}_{b,t}\}} \left(w^*_{a(k),b(k),t} - \mu_{b(k),t}\right) M_t \tilde{s}_{k,t}^{-a,b}(\tilde{\mathbf{p}}_t^{-a,b}; \theta^d)$$

$$d_{a,t}^{-a,b} = \sum_{k \in \mathscr{J}_{a,t} \setminus \mathscr{J}_{b,t}} \left(\tilde{p}_{k,t}^{-a,b} - w^*_{a,b(k),t} - c_{k,t}\right) M_t \tilde{s}_{k,t}^{-a,b}(\tilde{\mathbf{p}}_t^{-a,b}; \theta^d)$$

where $\tilde{\mathbf{p}}_t^{-a,b}$ denotes the J_t-dimensional vector of out-of-equilibrium retail prices when products purchased by the alliance a and sold under the brand b are no longer offered on market t, and $\tilde{s}_{k,t}^{-a,b}$ is the market share of each product k remaining on the market (see Section S4.2.2 of the Supplemental Material for further details).[52] From (11), I can derive the set of first-order conditions characterizing the surplus division of each bilateral negotiation in market t as follows:

$$\left(1 - \lambda_{f(j),a(j)}^{post}\right)\left(\pi_{f(j),t} - d_{f(j),t}^{-a(j),b(j)}\right) \frac{\partial \pi_{a(j),t}}{\partial w_{a(j),b(j),t}} + \lambda_{f(j),a(j)}^{post}\left(\pi_{a(j),t} - d_{a(j),t}^{-a(j),b(j)}\right) \frac{\partial \pi_{f(j),t}}{\partial w_{a(j),b(j),t}} = 0$$

$$\forall j \in \mathscr{J}_t \setminus \{0\}$$

By inverting this system of equations, I show in Section S4.2.2 of the Supplemental Material that one can obtain a closed-form expression for the J_t-dimensional vector of manufacturers' price-cost margins in the post-alliances periods, denoted by $\mathbf{\Gamma}_t^{post}$. The jth element of $\mathbf{\Gamma}_t^{post}$ is given by:

$$w_{a(j),b(j),t} - \mu_{b(j),t} = \Gamma_{a(j),b(j),t}^{post}(s_t, \mathbf{S}_{\mathbf{p}_t}, \mathbf{S}_{\mathbf{p}_t\mathbf{p}_t}, \tilde{\mathbf{S}}_{\Delta t}^{post}; \boldsymbol{\lambda}^{post}) \tag{12}$$

[52] The joint delisting decision of the alliance a implies that, in case of a bargaining breakdown, brand b is not distributed by any of its retailers.

where $\boldsymbol{\lambda}^{post}$ is a J_t-dimensional vector of manufacturers' bargaining weight with $\boldsymbol{\lambda}^{post}[j,1] = \lambda^{post}_{f(j),a(j)}$ and $\tilde{\mathbf{S}}^{post}_{\Delta t}$ is a $J_t \times J_t$ matrix of difference in market shares $s_{k,t}(\mathbf{p}_t;\theta^d) - \tilde{s}^{-a,b}_{k,t}(\tilde{\mathbf{p}}^{-a,b}_t;\theta^d)$.

In contrast to the price-cost margins of retailers which derive directly from estimates of demand parameters θ^d and the set of equations (7) characterizing necessary conditions for retailers' profit maximization, (10) and (12) show that the price-cost margins of manufacturers are identified up to two vectors of unknown parameters: $\boldsymbol{\lambda}^{pre}$ and $\boldsymbol{\lambda}^{post}$. In what follows, I describe my empirical strategy to estimate these parameters which allow to recover the price-cost margins of manufacturers and the division of surplus in the vertical chain before and after the formation of buyer alliances.

4.2.4 Identification and estimation of bargaining stage

To estimate $\boldsymbol{\lambda}^{pre}$ and $\boldsymbol{\lambda}^{post}$, I rely on an empirical approach that exploits the variation in the marginal costs of retailers across products and markets. Given that wholesale prices and distribution costs enter additively into the marginal costs of retailers, I make use of the following decomposition:

$$w_{j,t} + c_{j,t} = \underbrace{(w_{j,t} - \mu_{b(j),t})}_{\text{upstream market power}} + \underbrace{(\mu_{b(j),t} + c_{j,t})}_{\text{operational costs}} \quad \forall j,t$$

where the heterogeneity in the marginal costs of retailers is explained by differences in costs of production and distribution $(\mu_{b(j),t} + c_{j,t})$ and asymmetries in the ability of manufacturers to exert market power towards retailers $(w_{j,t} - \mu_{b(j),t})$. Without further assumptions, these two components are not separately identified as there are $\sum_{t=1}^{T} J_t$ equations and $2\sum_{t=1}^{T} J_t$ unknowns. I thus replace the upstream market power term by the closed-form expressions in (10) and (12) derived from the "Nash-in-Nash" bargaining solution. Without additional information on the marginal costs of products, I impose restrictions on their structure. In line with Gowrisankaran, Nevo and Town (2015), I assume that the constant marginal cost of production and distribution for product j in market t is given by:

$$\mu_{b(j),t} + c_{j,t} = \kappa_0 + \kappa_s x_{spark(j)} + \kappa_m x_{mineral(j)} + \kappa_{pl} x_{pl(j)} + \kappa_{b(j)} + \kappa_{r(j)} + \kappa_t + \omega_{j,t} \qquad (13)$$

where κ_0 denotes a constant term, $x_{spark(j)}$, $x_{mineral(j)}$, and $x_{pl(j)}$ are indicator variables for sparkling water, mineral water, and private label respectively, $\kappa_{b(j)}$, $\kappa_{r(j)}$, and κ_t are respectively national brand, retailer, and market fixed effects, and $\omega_{j,t}$ corresponds to an additive error term of unobserved cost factors (e.g., unobserved productivity).[53] This yields the following marginal cost function of retailers for product j in market t:

$$
\begin{aligned}
w_{j,t} + c_{j,t} = \quad & \Gamma^{pre}_{j,t}(\mathscr{s}_t, \mathbf{S}_{\mathbf{p}_t}, \mathbf{S}_{\mathbf{p}_t\mathbf{p}_t}, \tilde{\mathbf{S}}^{pre}_{\Delta t}; \boldsymbol{\lambda}^{pre}) \times \mathbb{1}\{\text{pre-alliances}\}_t \\
& + \Gamma^{post}_{a(j),b(j),t}(\mathscr{s}_t, \mathbf{S}_{\mathbf{p}_t}, \mathbf{S}_{\mathbf{p}_t\mathbf{p}_t}, \tilde{\mathbf{S}}^{post}_{\Delta t}; \boldsymbol{\lambda}^{post}) \times \mathbb{1}\{\text{post-alliances}\}_t \\
& + \kappa_0 + \kappa_s x_{spark(j)} + \kappa_m x_{mineral(j)} + \kappa_{pl} x_{pl(j)} + \kappa_{b(j)} + \kappa_{r(j)} + \kappa_t + \omega_{j,t}
\end{aligned}
\tag{14}
$$

where supply parameters to be estimated are the vectors of bargaining weights $\boldsymbol{\lambda}^{pre}$ and $\boldsymbol{\lambda}^{post}$, and the vector of cost parameters $\boldsymbol{\kappa} \equiv (\kappa_0, \kappa_s, \kappa_m, \kappa_{pl}, \boldsymbol{\kappa}_{\text{brand}}, \boldsymbol{\kappa}_{\text{retail}}, \boldsymbol{\kappa}_{\text{market}})$.[54] Note that the inclusion of $\omega_{j,t}$ enables a flexible specification of the marginal cost of products because the variation in $w_{j,t} + c_{j,t}$ that is left unexplained is attributed to unobserved cost factors.[55]

Identification assumptions. Given that the marginal costs of production and distribution are constant, it is worth noting that the bargaining weight parameters determine the slope of the marginal cost function of retailers. Hence, the variation in the market share of products should prove useful in separately identifying the contributions of the upstream market power and the operational costs to the retail marginal costs $w_{j,t} + c_{j,t}$. This source of variation, however, may not be exogenous. Indeed, whenever manufacturers and retailers observe marginal cost shocks before setting wholesale and retail prices, market shares are likely to be correlated with $\omega_{j,t}$. Market shares also depend on the unobserved quality $\xi_{j,t}$ which is likely correlated with unobserved cost factors in $\omega_{j,t}$. To address this concern, I adopt an instrumental variable approach that relies on three types of instruments affecting the price-cost margins of manufacturers (including market shares) but not the marginal costs of products.

The first set of instruments are the consumer characteristics, which shift demand but not marginal costs.[56] Intuitively, whenever manufacturers exert market power toward retailers (bargaining weights are different from 0), shifts in demand due to changes in consumer characteristics should explain variation in $w_{j,t} + c_{j,t}$. In practice, I use as instruments the proportion of consumers belonging to each of the four income groups in each market. To

[53]This constant marginal cost specification is common in empirical work on food industries (see, e.g., Villas-Boas, 2007; Miller and Weinberg, 2017; Michel, Paz y Miño and Weiergraeber, forthcoming).

[54]The vectors $\boldsymbol{\kappa}_{\text{brand}}$, $\boldsymbol{\kappa}_{\text{retail}}$, and $\boldsymbol{\kappa}_{\text{market}}$ include respectively all national brand, retailer, and market fixed effects.

[55]This relates to the cost specification in Gowrisankaran, Nevo and Town (2015). An alternative approach would be to specify the marginal costs of products in terms of only data and parameters, allowing the bargaining weights to depend on unobservables as in Grennan (2013).

[56]This source of variation is also commonly used to identify firm conduct in oligopolistic industries (e.g., Berry and Haile, 2014).

add flexibility, I interact these instruments with retailer fixed effects and a fixed effect for the post-alliances period. The validity of this first set of instruments hinges on the assumption that the characteristics of consumers are not correlated with unobserved cost shocks.

The second set of instruments exploits the amount of competition that each product faces in the characteristics space, which shifts the price-cost margins of manufacturers and should explain differences in market shares across products.[57] Intuitively, after controlling for differences in the marginal costs of production and distribution, products facing limited competition in the characteristics space should have systematically higher wholesale prices (and hence greater $w_{j,t} + c_{j,t}$) if manufacturers exert market power toward retailers. However, if the location of products in the characteristics space does not explain the differences in the marginal costs of retailers, this may provide information that retailers mitigate the market power of manufacturers ($w_{j,t}+c_{j,t} = \mu_{j,t}+c_{j,t}$). In practice, I use the number of products belonging to the same type of bottled water (e.g., mineral and sparkling) sold by the same retailer and by rival retailers. The validity of these two instrumental variables hinges on the assumption that observed product characteristics are exogenous (Berry, Levinsohn and Pakes, 1995).

The third set of instruments aligns with that used to identify consumer preferences for bottled water. More precisely, I interpret the formation of buyer alliances as an exogenous shifter of the bargaining environment that directly affects the price-cost margins of manufacturers. Based on Kalai (1977), the variation provided by the alliances should be particularly helpful to identify changes in firms' relative bargaining abilities (λ^{post}).[58] The logic is as follows. After controlling for common changes in the marginal costs of production and distribution before versus after the alliances, any remaining difference in the changes of $w_{j,t}+c_{j,t}$ between national brands purchased by an alliance and other products is interpreted as a shift in the bargaining weight of manufacturers vis-à-vis retailers involved in an alliance. In practice, I use the buyer alliances indicator variable $\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{inside alliances}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$ and the indicator variable for national brands that remain outside the scope of the alliances $\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$ as instrumental variables. Given the marginal cost specification in (13), the validity of these two instruments hinges on the assumption that changes in the unobserved cost factors of national brands purchased by an alliance, before versus after the formation of the alliances,

[57] This identification strategy builds on the *Differentiation IVs* developed by Gandhi and Houde (2023) to estimate consumer demand in differentiated products markets. See Gandhi and Houde (2020) and Michel, Paz y Miño and Weiergraeber (forthcoming) for a related use of such instruments to analyze firms' conduct in oligopolistic industries.

[58] In a similar spirit, Miller and Weinberg (2017) use the MillerCoors joint venture as an exogenous shifter of the competitive environment to identify changes in firms' conduct after the joint venture.

are not systematically different from changes in the unobserved cost factors of other products.[59] Because the effect of buyer alliances need not be uniform across products, I interact the buyer alliances indicator variable with manufacturer fixed effects and a fixed effect for each alliance. I also interact both indicator variables with each variable in the second set of instruments to capture the fact that the alliances may have affected products differently based on their location in the characteristics space. A complete list of the excluded instrumental variables is reported in Section S7.2 of the Supplemental Material (Table S3).

As the order condition requires at least one instrument for each bargaining weight (whose number grows with the number of manufacturer-retailer pairs) I impose the following parameter restrictions.[60] I consider that the bargaining weights of manufacturers when dealing with retailers 7 and 8 are identical and do not vary over time. For retailers involved in an alliance, I allow the bargaining weights to differ before and after the alliances. In the pre-alliances periods, I use a size-based restriction by considering that manufacturers have identical bargaining weights when negotiating with large retailers (retailers 1, 3, and 5) and with small retailers (retailers 2, 4, and 6) according to Table 1. In the post-alliances periods, the bargaining weights are assumed to be alliance-specific (i.e., manufacturers have identical bargaining weights with respect to each buyer alliance).

Estimation procedure. Supply parameters $\boldsymbol{\lambda}^{pre}$, $\boldsymbol{\lambda}^{post}$, and $\boldsymbol{\kappa}$ in (14) are estimated by two-step GMM (Hansen, 1982). In practice, I concentrate the cost parameters $\boldsymbol{\kappa}$ out of the GMM objective function and search nonlinearly over the vector of bargaining weights $\boldsymbol{\lambda} \equiv (\boldsymbol{\lambda}^{pre}, \boldsymbol{\lambda}^{post})$. Formally, the resulting GMM estimator is given by:

$$\hat{\boldsymbol{\lambda}} \equiv \underset{\boldsymbol{\lambda}}{\operatorname{argmin}} \, (\mathbf{Z}^s \boldsymbol{\omega}(\boldsymbol{\lambda}, \boldsymbol{\kappa}(\boldsymbol{\lambda})))^\top \mathbf{A}^{-1} \mathbf{Z}^s \boldsymbol{\omega}(\boldsymbol{\lambda}, \boldsymbol{\kappa}(\boldsymbol{\lambda})) \tag{15}$$

where $\boldsymbol{\omega}(\boldsymbol{\lambda}, \boldsymbol{\kappa}(\boldsymbol{\lambda}))$ is the $\sum_t J_t$-dimensional vector of unobserved cost factors, $\mathbf{Z}^s$ is a $K \times \sum_t J_t$ matrix of instrumental variables, and $\mathbf{A}$ is a $K \times K$ weighting matrix (see Section 7.1 of the Supplemental Material for a detailed description of the estimation procedure).

[59]This assumption would be violated if, for instance, the alliances have generated cost efficiencies. For reasons stated at the beginning of Section 4.2 (remark on the contractual form) I argue that such efficiencies are unlikely to arise.

[60]I have 21 manufacturer-retailer pairs in the pre-alliances periods and 15 manufacturer-alliance pairs in the post-alliances periods, resulting in a total of 36 bargaining weights. As it is hard to find enough good instruments to identify each bargaining weight in practice, I follow prior empirical work by imposing restrictions to reduce the number of parameters to be estimated (see also Gowrisankaran, Nevo and Town, 2015; Ho and Lee, 2017; Crawford et al., 2018).

Table 2: Estimates of consumer demand

(a) **Consumer taste heterogeneity**

Variable	Coef.	S.E.
Observed heterogeneity (σ^o):		
Price × income (0–900)	ref.	
Price × income (900–1,899)	−0.18*	0.08
Price × income (1,900–4,449)	−0.47*	0.08
Price × income (>4,449)	−0.79*	0.08
Mineral × age (18–40)	ref.	
Mineral × age (40–60)	0.54*	0.01
Mineral × age (>60)	1.04*	0.01
Unobserved heterogeneity (σ^u):		
Constant	1.84*	0.01
Price	3.49*	0.02
Mineral	1.73*	0.01
Sparkling	2.13*	0.01
Private label	1.86*	0.01
Convergence criterion ($g^{\top}H^{-1}g$)	6.55×10^{-6}	
Simulated log-likelihood	$-1,840,320$	
Number of observations	550,059	

Notes: Simulated maximum likelihood estimates using 200 shuffled Halton draws and the BHHH method. "Price" is the retail price variable, and the monthly income intervals are in euros. * indicates significance at the 5% level. The convergence criterion is based on the statistic $g^{\top}H^{-1}g$, where g and H denote respectively the gradient and the hessian matrix.

(b) **Mean taste parameters**

	OLS		TSLS	
Variable	Coef.	S.E.	Coef.	S.E.
Price	−5.90*	0.40	−17.54*	3.21
Mineral	−0.20	0.11	0.54*	0.24
Sparkling	−1.68*	0.16	−1.09	0.21
Constant	−4.21*	0.18	−2.45*	0.52
Brand fixed effects:				
Brand 1	2.87*	0.16	7.27*	1.22
Brand 2	−0.03	0.11	0.76*	0.24
Brand 3	0.97*	0.13	3.77*	0.78
Brand 4	0.05	0.09	0.39*	0.14
Brand 5	1.65*	0.16	5.98*	1.20
Brand 6	0.50*	0.10	1.99*	0.43
Brand 7	−0.12	0.09	0.47*	0.20
Brand 8	1.17*	0.08	1.91*	0.22
Brand 9	2.08*	0.15	6.05*	1.11
Brand 10	0.44*	0.13	−0.46	0.27
Brand 11	−0.89*	0.10	−0.74*	0.12
Private label	ref.		ref.	
Retailer fixed effects:				
Retailer 1	1.81*	0.16	2.45*	0.22
Retailer 2	−0.19	0.16	0.62*	0.26
Retailer 3	1.22*	0.16	2.03*	0.25
Retailer 4	0.80*	0.16	1.44*	0.22
Retailer 5	1.34*	0.16	1.98*	0.21
Retailer 6	0.78*	0.16	1.60*	0.26
Retailer 7	1.76*	0.16	2.31*	0.19
Retailer 8	ref.		ref.	
Market fixed effects (not shown)				
Standard F			23.35	
F_{eff}			20.73	
Number of observations	2,192			

Notes: "Price" stands for the retail price variable. Heteroskedasticity-robust standard errors. * indicates significance at the 5% level. Standard F indicates the standard first-stage F-statistic and F_{eff} shows the heteroskedasticity-robust first-stage F-statistic from Montiel Olea and Pflueger (2013) for the TSLS estimates.

4.3 Estimation results

4.3.1 Consumer demand estimates

Table 2 presents the coefficient estimates from the demand model described in Section 4.1. Estimates of the mean taste parameters are reported in Table 2b. Consistent with Berry (1994) and related literature, the OLS underestimate the retail price sensitivity compared to the TSLS. I thus focus on the TSLS estimates for the remainder of the article. The robust first-stage F-statistic suggested by Montiel Olea and Pflueger (2013) is well above the critical

Table 3: Estimates of substitution patterns

	Own-price elasticity	Cross-price elasticity		Recapture ratio	
		Within	Outside	Within	Outside
Inside goods	−7.07	0.05	0.01	83.65	16.35
National brands	−8.01	0.07	0.03	67.18	32.82
Private labels	−4.44	0.07	0.02	53.66	46.34
Spring water	−3.17	0.06	0.02	46.76	53.24
Mineral water	−7.99	0.06	0.02	67.17	32.83
Still water	−5.31	0.05	0.03	64.78	35.22
Sparkling water	−9.92	0.13	0.03	51.62	48.38

Notes: Average own- and cross-price elasticities of demand and recapture ratios. Following a 1% retail price increase of a product in the category of row j (e.g., mineral water), the average cross-price elasticities indicates the percentage change in market share for another product within and outside that category. The recapture ratios (also known as "aggregate diversion ratios") measure the average fraction of sales lost by a product in the category of row j (e.g., mineral water) following a retail price increase that is recaptured by products within and outside that category.

value of 6.36, indicating that the instrumental variables are not weakly correlated with the endogenous retail price variable (see Table 11 of Appendix C for further details on the first stage regression). Most of the parameters are precisely estimated and have the expected sign. In line with Bonnet and Dubois (2015), I find that households have a positive mean valuation for mineral water. Brand fixed effects also show an important heterogeneity in the mean valuation across brands of bottled water. Table 2a presents estimates of demand parameters that capture the heterogeneity in consumer preferences. The convergence criterion $g^{\top}H^{-1}g$, where g and H are respectively the gradient and the Hessian matrix, takes the value 6.55×10^{-6}. This indicates that all elements of the gradient are very close to 0 (see Train, 2009). The σ^o parameters capturing the degree of observed heterogeneity in consumer preferences are accurately estimated and reveal an intuitive pattern: richer households exhibit lower price sensitivity and older households have a higher valuation for mineral water. Consistent with Berry, Levinsohn and Pakes (2004), such parameters do not explain the full pattern of substitution in the data. Indeed, the σ^u parameters measuring the degree of unobserved heterogeneity in consumer preferences are fairly large and precisely estimated.[61]

Using demand estimates, I compute the own and cross-price elasticity of demand. Table 3 reports the results for different categories of bottled water. The average own-price elasticity equals −7.07. There is however substantial heterogeneity across product categories. Consistent with Bonnet and Dubois (2010, 2015), the own-price elasticity for mineral waters is on average more than twice as high as that for spring waters.[62] The table also

[61] The estimated normal distribution for the retail price sensitivity across consumers has a mean of −17.54 and a standard deviation of 3.49, implying that essentially no consumer exhibits positive price sensitivity.

[62] The estimated average own-price elasticity for spring waters is −3.17, which is in line with estimates in Bonnet and Dubois (2015) and Zhang (2018).

reports the average cross-price elasticity between products of the same category (within) and other categories (outside). For instance, a 1% price increase in a mineral water product increases on average the market share of another mineral water product by 0.06% and that of a non-mineral water product by 0.02%. Again, these estimates are in line with Bonnet and Dubois (2010). To gain additional insight into the estimated substitution patterns, the table also displays the recapture ratios (also known as "aggregate diversion ratios"). These ratios measure the average fraction of sales lost by a product due to a retail price increase that is recaptured by all products of the same category (within) and other categories (outside). For instance, 67.17% of the demand lost by a mineral water product due to a retail price increase diverts to other mineral water products. This demand diversion is much lower for non-mineral water products (32.83%). Overall, these results reveal reasonable and intuitive substitution patterns among products.

4.3.2 Downstream competition and bargaining estimates

Table 4 presents estimates of the marginal cost function of retailers described in (14). Note that although the bargaining weight parameters should lie in the interval $[0, 1]$, no constraints are imposed on these parameters during estimation. The value of the GMM objective function at the estimated parameters is 52.99 and the Hansen's (1982) overidentification test cannot reject the model at the 95% confidence level. Moreover, the battery of diagnostic tests for weak instruments conducted in Section S7.2 of the Supplemental Material provides evidence regarding the strength of the excluded instruments used to identify the bargaining weight parameters.[63]

Cost parameters are precisely estimated and have the expected signs (e.g., both mineral and sparkling fixed effects contribute positively to the marginal costs of products). In the pre-alliances periods, the estimated bargaining weights of manufacturers relative to retailers involved in an alliance are above 0.50. Moreover, manufacturers have a greater bargaining weight vis-à-vis small retailers. This result is in line with bargaining estimates of Crawford and Yurukoglu (2012) who find that small cable operators and satellite providers have lower bargaining weights than large cable operators in the multichannel television industry. In the post-alliances periods, estimates indicate that the bargaining weights of manufacturers vis-à-vis retailers involved in an alliance have been reduced. As described in Section 2, this is consistent with Kalai's (1977) theory according to which a coalition of bargainers should be treated more favorably by the Nash solution through asymmetric weights. This finding can

[63] Following Gandhi and Houde (2023), these first-stage diagnostic tests are based on a Gauss-Newton regression model.

Table 4: Bargaining and marginal cost parameter estimates

Variable	Coefficient	S.E.
Cost parameters:		
Constant	0.10*	0.01
Mineral	0.06*	0.00
Sparkling	0.05*	0.00
Brand fixed effects (not shown)		
Retailer fixed effects (not shown)		
Market fixed effects (not shown)		
Bargaining weights (pre-alliances):		
Large retailers (retailers 1, 3, 5)	0.55	0.20
Small retailers (retailers 2, 4, 6)	0.74	0.21
Bargaining weights (post-alliances):		
Buyer alliance 1 (retailers 1, 2)	0.02⋆	0.17
Buyer alliance 2 (retailers 3, 4)	0.36	0.17
Buyer alliance 3 (retailers 5, 6)	0.13⋆	0.17
Retailers 7, 8	0.00	–
GMM objective function value	52.99	
χ^2 critical value (5%)	66.34	
Number of observations	2,192	

Notes: 2-step GMM estimates. Heteroskedasticity-robust standard errors. * indicates significance at the 5% level. ⋆ indicates bargaining weight estimates that differ significantly from 0.5 at the 5% level.

also be related to Lewis and Pflum (2015) who provide evidence that hospital systems have higher bargaining weights vis-à-vis insurers than individual hospitals.

I find an estimated value of 0 for the bargaining weight of manufacturers vis-à-vis retailers 7 and 8, suggesting that the latter make take-it-or-leave-it offers to manufacturers.[64] Although the analysis of alliance formation is inherently complex and beyond the scope of my article, these bargaining estimates are particularly consistent with the fact that retailers 7 and 8 have not formed any buyer alliance (see Figure 3 for an illustrative example). Based on these parameter estimates, I compute the price-cost margins of manufacturers, the marginal costs of products, and the surplus division between manufacturers and retailers before and after the formation of buyer alliances. The results are reported in Table 5.

In the pre-alliances periods, the table shows that retailers involved in an alliance have a quantity-weighted average price-cost margin of 26.56%, which is lower than that of retailers 7 and 8 equal to 37.56%. These values slightly increase in the post-alliances periods. On the contrary, the price-cost margins of manufacturers fall from 13.60% to 4.09%, which is mainly due to the substantial decline in their margins on products sold by retailers involved in a buyer alliance (see the first two columns). The last row of the table reports the average

[64] I do not report any standard error for this estimated bargaining weight because it lies on the boundary of the parameter space. Instead, I use a grid search method and select the value of the bargaining weight that minimizes the GMM objective function (see Section S7.1 of the Supplemental Material for further details).

Table 5: Price-cost margins, marginal costs, and surplus division

	Inside alliances		Outside alliances		Total	
	Pre	Post	Pre	Post	Pre	Post
Price-cost margins:						
Retailers: $\boldsymbol{\gamma}$	26.56	27.45	37.56	39.08	30.16	31.22
	(0.47)	(0.53)	(0.93)	(1.12)	(0.52)	(0.69)
Manufacturers: $\boldsymbol{\Gamma}$	17.05	5.04	3.92	1.25	13.60	4.09
	(0.49)	(0.24)	(0.21)	(0.10)	(0.43)	(0.20)
Marginal cost:						
Retailers: $\mathbf{w}+\mathbf{c}$	0.28	0.26	0.20	0.19	0.25	0.24
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Industry: $\mathbf{c}+\boldsymbol{\mu}$	0.25	0.25	0.20	0.18	0.23	0.23
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Division of surplus:						
Retailers' share	59.43	84.45	93.59	98.07	65.19	87.25
	(0.21)	(0.44)	(0.33)	(0.07)	(0.38)	(0.41)

Notes: Standard deviations in parenthesis refer to variation across markets. Terms "Pre" and "Post" stand respectively for the pre-alliances periods and the post-alliances periods. The column labeled "Inside alliances" presents the results for retailers that have formed a buyer alliance. The column labeled "Outside alliances" presents the results for retailers 7 and 8. Average price-cost margins in percentage of retail prices and average marginal costs in euros per liter are calculated using quantity weights. The last row corresponds to the average share of the surplus generated from bilateral agreements captured by retailers (in percentage).

share that retailers capture from the surplus generated by bilateral agreements with manufacturers. In the pre-alliances periods, results show that retailers involved in an alliance capture on average 59.43% of this surplus, implying that they have a higher bargaining power than manufacturers. The second column indicates that this bargaining power is reinforced in the post-alliances periods, where retailers capture 84.45% of the surplus resulting from bilateral agreements.

5 Counterfactual analysis

In this section, I use the estimated structural parameters to study the mechanisms through which a buyer alliance affects market outcomes in bilateral oligopolies. Following the theoretical insights of Section 2, I examine three counterfactual scenarios: (a) the buyer alliances do not occur (no alliance), (b) the buyer alliances occur but only with a status quo effect (joint delisting decision), and (c) the buyer alliances occur without affecting the bargaining ability of manufacturers vis-à-vis retailers (joint-bargaining paradox). Under each scenario, I compute a new market equilibrium in negotiated wholesale prices, retail prices, and market shares of products during the post-alliances periods by holding fixed the marginal costs of production and distribution, consumer preferences, product characteristics, and the manufacturer-retailer network structure. The iterative algorithm used to perform each

Table 6: Results of the counterfactual scenarios

	%∆ Retail price	%∆ Margins		%∆ Profit			%∆ Welfare	
		Retail.	Manuf.	Retail.	Manuf.	Industry	Cons.	Total
Scenario A: No Alliance								
Inside alliances	−10.01	3.01	−72.58	31.23	−64.67	7.72	–	–
	(0.19)	(0.10)	(0.79)	(0.59)	(0.30)	(0.29)		
Outside alliances	−1.03	−1.01	−67.18	−10.18	−70.05	−12.36	–	–
	(0.07)	(0.09)	(0.92)	(0.17)	(0.95)	(0.19)		
Total	−5.37	0.93	−71.22	5.43	−65.12	−10.06	8.39	2.09
	(0.15)	(0.06)	(0.83)	(0.17)	(0.95)	(0.19)	(0.19)	(0.33)
Scenario B: Joint delisting decision								
Inside alliances	−3.21	0.46	−21.91	9.07	−18.50	−2.23	–	–
	(0.07)	(0.06)	(0.62)	(0.23)	(0.31)	(0.12)		
Outside alliances	−0.35	−0.43	−20.70	−3.51	−23.12	−4.22	–	–
	(0.02)	(0.03)	(0.33)	(0.07)	(0.45)	(0.08)		
Total	−1.73	0.00	−21.53	1.23	−18.89	−3.19	2.69	0.68
	(0.06)	(0.04)	(0.48)	(0.07)	(0.31)	(0.07)	(0.06)	(0.11)
Scenario C: Joint-bargaining paradox								
Inside alliances	−1.18	0.15	−7.37	2.97	−5.16	−0.34	–	–
	(0.03)	(0.02)	(0.28)	(0.14)	(0.25)	(0.06)		
Outside alliances	−0.11	−0.16	−6.05	−1.23	−7.14	−1.44	–	–
	(0.01)	(0.01)	(0.27)	(0.04)	(0.35)	(0.05)		
Total	−0.63	−0.01	−7.04	0.35	−5.33	−0.89	0.90	0.28
	(0.02)	(0.01)	(0.28)	(0.02)	(0.26)	(0.05)	(0.03)	(0.03)

Notes: Rows "Inside alliances" and "Outside alliances" refer respectively to products within the scope of the alliances (national brands sold by retailers 1 to 6) and outside the scope of the alliances (national brands sold by retailers 7 and 8 and all private labels). Standard deviations in parenthesis refer to variation across markets. Percentage changes in retail prices and price-cost margins are calculated using quantity weights.

counterfactual experiment is described in Section S9 of the Supplemental Material and the results are reported in Table 6.

Scenario A: No Alliance. This counterfactual scenario simulates equilibrium market outcomes in the absence of buyer alliances to analyze their effects on both wholesale and retail prices. Results in Table 6 show that the baseline quantity-weighted price-cost margins of manufacturers over national brands purchased by an alliance are 72.58% lower relative to the "No Alliance" scenario. Furthermore, by deteriorating the status quo position of manufacturers in their bargaining with retailers 7 and 8, I find that the alliances also reduce the quantity-weighted price-cost margins of manufacturers over national brands remaining outside their scope by 67.18%.[65] This downward shift in the price-cost margins of manufacturers (and hence wholesale prices) is passed on to consumers in the form of a 5.37% decrease in the quantity-weighted average retail price. Therefore, simulation results suggest

[65]The status quo position of manufacturers in their bargaining with retailers 7 and 8 corresponds to the profit they get from bilateral agreements with each buyer alliance. Hence, by decreasing this profit, the alliances reinforce the bargaining power of retailers 7 and 8 vis-à-vis manufacturers.

that the formation of buyer alliances has generated a countervailing buyer power effect in line with Galbraith's (1952) view. Indeed, retailers involved in an alliance and consumers benefit from this countervailing force to the detriment of other retailers, manufacturers, and the industry which respectively experience a profit decrease of 10.18%, 65.12%, and 10.06% relative to the "No Alliance" scenario.[66] These findings are consistent with recent political debates about the effect of retailers' buyer power on the surplus generated by food industries and its division within the supply chain.[67]

A comparison of the retail price effect of the buyer alliances obtained under the counterfactual simulation and the event-study analysis in Section 3.3 can provide some insights into the robustness of my estimates. First, the simulation results indicate that the unweighted average retail price of private labels remains almost unaffected by the formation of buyer alliances (+0.04%), which provides support for the identification assumption in the event study. Second, the simulation suggests that the alliances caused a decline of 8.15% in the unweighted average retail price of national brands purchased by an alliance, which aligns fairly well with the retail price decrease of 7.23% estimated in the event-study analysis. Third, the simulation predicts that the unweighted average retail price of national brands sold by retailers 7 and 8 decreases by 2.33%, which also closely aligns with the 2.50% retail price decrease estimated in the event-study analysis (although this point estimate is not significant at the 5% level). Therefore, it is reassuring that the counterfactual simulation and the event-study analysis provide qualitatively similar results regarding the average retail price effect of the buyer alliances.

To my knowledge, this counterfactual simulation provides the first empirical evidence of a change in market structure that creates countervailing buyer power as envisioned by Galbraith (1952). In particular, this contrasts with the countervailing force identified in Ho and Lee (2017) which arises to the detriment of downstream firms (insurers) and only when downstream prices (premiums) are determined through negotiations with consumers (employer).[68] Based on the theoretical insights developed in Section 2, I rely on two other counterfactual scenarios to analyze the determinants of countervailing buyer power.

[66] I use the log-sum formula to compute the change in consumer surplus (e.g., Train, 2009, Chapter 3). Given that the utility specification (2) is a piece-wise linear spline function of income, this provides an accurate approximation of the true change in consumer surplus which does not admit any closed-form expression (see Morey, Sharma and Karlstrom, 2003).

[67] For instance, a National Food Conference involving all stakeholders (e.g., agrifood companies, food retailers, consumer organizations) was organized by the French government in 2017 to revitalize the surplus generated by food industries and achieve a fairer division of this surplus along the supply chain.

[68] It also differs from Gowrisankaran, Nevo and Town (2015) who rely on an ex-ante analysis of hospital mergers to reveal the presence (but not the creation) of countervailing power exerted by insurers.

Scenario B: Joint delisting decision. To isolate the status quo effect of buyer alliances described in Section 2, I simulate the formation of buyer alliances in which retailers keep negotiating wholesale prices separately and secretly and do not benefit from a bargaining ability effect (that is, the bargaining weights of manufacturers vis-à-vis retailers remain similar to the pre-alliances periods). However, if a retailer involved in an alliance fails to reach an agreement with a manufacturer on one of its national brands, I consider that all retailers in the alliance jointly delist the manufacturer's brand from their shelves (see Section S4.2.3 of the Supplemental Material for further details). The results reported in Table 6 are interpreted relative to the "No Alliance" scenario. I find that the status quo effect of buyer alliances generates a decrease of 21.53% in the quantity-weighted price-cost margins of manufacturers relative to the situation without the alliances. This decline in wholesale prices is passed on to retail prices, with a 1.73% decrease in the quantity-weighted retail prices. Simulation results also indicate that the profits of retailers involved in an alliance increase by 9.07% to the detriment of other retailers, manufacturers, and the industry profit. Table 13 of Appendix D provides additional results for each retailer. In line with the theoretical insights from Section 2, this table shows that every retailer of a buyer alliance obtains lower wholesale prices relative to the "No Alliance" scenario. Interestingly, however, I find that retailer 1's profit decreases by 3.60%. As illustrated in Figure 3, this suggests that the gain in the status quo position of retailer 1 (the largest retailer) from forming an alliance with retailer 2 (the smallest retailer) is not enough to compensate for the losses of improving the competitiveness of retailer 2 in the downstream market. Consequently, an alliance with retailer 2 is never profitable for retailer 1, except if it benefits from a bargaining ability effect. In the last counterfactual scenario, I further explore the role of this bargaining effect on the countervailing force and the profitability of buyer alliances.

Scenario C: Joint-bargaining paradox. This counterfactual scenario simulates the formation of buyer alliances that do not affect the bargaining ability of firms (that is, the bargaining weights of manufacturers vis-à-vis retailers remain similar to the pre-alliances periods). Consistent with the theoretical predictions from Section 2, simulation results reported in Table 6 indicate that the countervailing force generated by the status quo effect of the alliances is substantially reduced by the nondiscrimination effect. While the quantity-weighted price-cost margins of manufacturers fall by 7.04%, the quantity-weighted retail prices decrease by only 0.63% relative to the "No Alliance" scenario.

The results presented in Table 14 of Appendix D are even more striking. For each buyer alliance, I find that the largest retailer systematically gets lower profits relative to the situation when it bargains alone. As illustrated by the gray areas of Figure 3, this result is reminiscent of the Harsanyi's (1977) joint-bargaining paradox according to which a coalition of multiple bargainers who act in concert during negotiations obtains less favorable trading terms than what each bargainer can negotiate on his own. The unprofitability of group purchasing in bilateral oligopolies mainly stems from the nondiscrimination effect. First, this effect makes the purchasing conditions of retailers more uniform, thereby depriving some of them of a competitive advantage in the downstream market. Second, it also weakens the bargaining power of retailers vis-à-vis manufacturers which, in turn, undermines the countervailing force stemming from the status quo effect. Therefore, absent any increase in the bargaining ability of retailers, I find that the buyer alliances observed in the French food retail sector in 2014 would not have been formed.[69] This result has important implications for the analysis of market structure changes. In particular, it provides direct evidence that a complete understanding of the effects of any market concentration on bargaining outcomes requires analyzing the impact on bargaining abilities, which has been largely ignored in the empirical literature.

6 Concluding remarks

Leveraging a unique case in the French food retail sector in which large retailers joined forces to negotiate common wholesale prices with manufacturers through the formation of three buyer alliances, this article studies the determinants of buyer power and their effects on market outcomes. I shed light on three main economic forces: (i) a status quo effect which strengthens the bargaining position of retailers vis-à-vis manufacturers, (ii) a nondiscrimination effect that alters firms' relative costs of making price concessions to the benefit of manufacturers, and (iii) a bargaining ability effect that enhances the bargaining power of retailers. Using pre- and post-alliances data on household purchases of bottled water, I develop a structural model of bilateral oligopoly that allows to directly estimate the buyer alliances effect on each source of bargaining power in the vertical chain.

The results indicate that the bargaining power of retailers is higher than that of manufacturers and has been reinforced after the formation of buyer alliances. In a set of counterfactual experiments, I provide evidence of a countervailing buyer power effect that generates a retail price decrease of 8.15% for national brands purchased by a buyer alliance. Moreover, I find that retailers' asymmetries and changes in bargaining abilities play an important role

[69] In an oligopoly framework with negotiated prices, Grennan (2013) derives similar conclusions regarding the profitability of a group purchasing organization formed by hospitals operating on different local markets.

in the emergence of this countervailing force and the profitability of buyer alliances.

This article is subject to a certain number of limitations that I leave for future research. First, my empirical results suggest that the impact of changes in the bargaining ability of firms quantitatively matters in the analysis of market concentration. Although my approach has the advantage of not imposing any particular structure on this source of bargaining power, a better understanding of its determinants constitutes an important research agenda (e.g., Grennan, 2014). Second, by focusing on the bottled water industry, this article offers a partial analysis of the formation of buyer alliances which covered other product categories as well (e.g., dairy products). A possible direction for future work is to combine the bilateral oligopoly framework developed in this article with a model of cross-category pricing (e.g., Thomassen et al., 2017) to study the implications of multi-category effects on buyer power. Finally, this article highlights that the countervailing force of buyer alliances arises to the detriment of both manufacturers and the industry surplus, which is likely to have long-run consequences. How the exercise of buyer power affects product variety and innovation raises important concerns for policymakers (European Commission, 2014). The lack of clear theoretical predictions on this topic (e.g., Montez, 2007; Caprice and Rey, 2015; Chen, 2019) offers an exciting avenue for empirical work.

Appendix

A Additional descriptive statistics

Table 7: Descriptive statistics for brands

	Mineral	Sparkling	Market share		Retail price	
			pre-alliances	post-alliances	pre-alliances	post-alliances
National brands:						
Brand 1	Yes	Yes	4.38	4.66	0.72	0.68
			(0.64)	(0.61)	(0.02)	(0.03)
Brand 2	Yes	No	3.43	3.28	0.36	0.32
			(0.28)	(0.40)	(0.02)	(0.01)
Brand 3	Yes	No	3.21	3.89	0.53	0.50
			(0.30)	(0.32)	(0.01)	(0.01)
Brand 4	Yes	No	3.24	3.40	0.31	0.28
			(0.34)	(0.34)	(0.02)	(0.01)
Brand 5	Yes	Yes	1.43	1.49	0.75	0.65
			(0.35)	(0.19)	(0.03)	(0.01)
Brand 6	Yes	No	3.46	3.09	0.41	0.40
			(0.24)	(0.23)	(0.02)	(0.02)
Brand 7	Yes	No	2.52	2.58	0.33	0.33
			(0.32)	(0.19)	(0.01)	(0.01)
Brand 8	Yes	Yes	2.53	2.68	0.41	0.38
			(0.18)	(0.13)	(0.01)	(0.00)
Brand 9	Yes	Yes	2.35	2.22	0.70	0.64
			(0.30)	(0.20)	(0.03)	(0.05)
Brand 10	No	No	11.62	13.19	0.13	0.13
			(0.67)	(0.86)	(0.00)	(0.00)
Brand 11	Yes	No	1.47	1.52	0.31	0.27
			(0.15)	(0.18)	(0.00)	(0.01)
Private labels:						
PL 1	No	No	13.87	12.89	0.20	0.21
			(0.52)	(0.36)	(0.01)	(0.01)
PL 2	No	Yes	0.66	0.43	0.25	0.26
			(0.06)	(0.03)	(0.01)	(0.02)
PL 3	Yes	No	3.53	3.47	0.27	0.27
			(0.19)	(0.11)	(0.01)	(0.01)
PL 4	Yes	Yes	5.34	4.53	0.32	0.32
			(0.26)	(0.24)	(0.02)	(0.01)

Notes: Market shares in percentage are calculated according to the number of household purchases. Average retail prices (in euro per liter). Standard deviation depicts variation across markets. Brand 1 to 5 are produced by manufacturer 1, brand 6 to 9 are produced by manufacturer 2, and brand 10 to 11 are produced by manufacturer 3. Remark that I am not permitted to reveal names of brands due to confidentiality regarding Kantar WorldPanel data.

Table 8: Descriptive statistics for national brands and private labels of retailers

	Market share				Retail price			
	national brand		private label		national brand		private label	
	pre	post	pre	post	pre	post	pre	post
Retailer 1	11.27	11.45	3.58	2.76	0.46	0.41	0.26	0.26
	(0.33)	(0.46)	(0.22)	(0.25)	(0.01)	(0.02)	(0.01)	(0.01)
Retailer 2	1.39	1.50	0.39	0.36	0.46	0.41	0.30	0.31
	(0.15)	(0.16)	(0.05)	(0.05)	(0.02)	(0.03)	(0.02)	(0.03)
Retailer 3	5.94	6.06	1.37	1.15	0.46	0.43	0.28	0.29
	(0.38)	(0.47)	(0.10)	(0.08)	(0.02)	(0.02)	(0.01)	(0.01)
Retailer 4	3.21	4.35	1.74	1.64	0.45	0.42	0.26	0.25
	(0.24)	(0.15)	(0.09)	(0.07)	(0.01)	(0.01)	(0.01)	(0.01)
Retailer 5	4.43	4.78	4.54	4.16	0.45	0.41	0.28	0.26
	(0.70)	(0.33)	(0.17)	(0.23)	(0.01)	(0.02)	(0.01)	(0.01)
Retailer 6	3.00	3.27	1.62	1.48	0.47	0.43	0.29	0.27
	(0.20)	(0.13)	(0.08)	(0.20)	(0.01)	(0.01)	(0.01)	(0.01)
Retailer 7	9.65	9.94	4.83	4.23	0.44	0.40	0.25	0.29
	(0.57)	(0.46)	(0.31)	(0.11)	(0.01)	(0.01)	(0.00)	(0.03)
Retailer 8	0.76	0.65	5.34	5.54	0.33	0.32	0.20	0.20
	(0.09)	(0.09)	(0.13)	(0.25)	(0.02)	(0.01)	(0.01)	(0.02)

Notes: Terms “pre” and “post” stand respectively for the pre-alliances periods (2013) and the post-alliances periods (2015). Market shares in percentage are calculated according to the number of household purchases. Average retail prices are in euro per liter. Standard deviations in parenthesis refer to variation across markets for the year 2013 (pre-alliances) and 2015 (post-alliances). Remark that I am not permitted to reveal names of retailers due to confidentiality regarding Kantar WorldPanel data.

Table 9: **Descriptive statistics for household characteristics**

	Percentage		Retail price		Mineral	
	pre	post	pre	post	pre	post
Household income range:						
0–900	2.21	1.97	0.29	0.31	52.66	59.21
	(0.09)	(0.11)	(0.01)	(0.01)	(2.43)	(2.29)
900–1,899	22.28	21.65	0.30	0.29	56.89	55.68
	(0.54)	(0.57)	(0.01)	(0.01)	(1.14)	(0.96)
1,900–4,449	54.30	54.61	0.32	0.30	57.67	57.38
	(0.51)	(0.39)	(0.01)	(0.01)	(1.37)	(1.08)
>4,500	21.20	21.77	0.34	0.32	63.37	62.72
	(0.67)	(0.53)	(0.01)	(0.01)	(1.08)	(1.03)
Age of household head:						
18–40	27.96	27.44	0.29	0.28	50.05	49.78
	(1.04)	(0.85)	(0.01)	(0.01)	(1.43)	(1.60)
40–60	46.03	46.23	0.32	0.30	58.74	58.16
	(0.50)	(0.77)	(0.01)	(0.01)	(1.14)	(0.74)
>60	26.01	26.33	0.34	0.32	66.01	65.26
	(1.20)	(1.32)	(0.01)	(0.01)	(1.00)	(0.64)
All households:						
	100	100	0.32	0.30	58.55	58.19
	–	–	(0.01)	(0.01)	(1.09)	(0.89)
Average number of households per market: 11,066.98						

Notes: Terms "pre" and "post" stand respectively for the pre-alliances periods (2013) and the post-alliances periods (2015). Household income range in euro per month. Column "Percentage" reports the percentage of each household group in the database. Column "Retail price" reports the average retail price (in euro per liter) of products purchased by each household group. Column "Mineral" shows the percentage of households purchasing mineral water for each household group. Standard deviations in parenthesis refer to variation across markets for the year 2013 (pre-alliances) and 2015 (post-alliances).

B Event study analysis: Robustness to alternative specifications

In this section, I explore the robustness of the estimates obtained from the event-study analysis developed in Section 3.3.

B.1 Pre- and post-alliances aggregation

An alternative solution for addressing the serial correlation problem in the regression model (1) compared to clustering standard errors is to aggregate the data before and after the formation of buyer alliances (Bertrand, Duflo and Mullainathan, 2004). Based on this insight, I average the data pre- and post-alliances (panel of length 2) and consider the following regression model:

$$\begin{aligned} \overline{\ln p}_{j,t} = \; & \beta^{in}_{post} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t \\ & + \beta^{out}_{post} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t \\ & + \beta_j + \beta_{post} \mathbb{1}\{\text{post-alliances}\}_t + \overline{u}_{j,t} \end{aligned} \quad (16)$$

Table 10: Results of the pre- and post-alliances aggregation

Variable	Coefficient	S.E.
Dependent variable: $\overline{\ln p}_{j,t}$		
$\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$	-0.076^*	0.0215
$\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$	-0.059^*	0.0242
$\mathbb{1}\{\text{post-alliances}\}_t$	-0.006	0.0192
Product fixed effects (not shown)		
R^2 adjusted	0.998	
Number of observations	220	
Dependent variable: $\overline{\ln p^{detrend}_{j,t}}$		
$\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$	-0.076^*	0.0215
$\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$	-0.017	0.0242
$\mathbb{1}\{\text{post-alliances}\}_t$	-0.006	0.0192
Product fixed effects (not shown)		
R^2 adjusted	0.998	
Number of observations	220	

Notes: OLS estimates of the regression model (16) with either $\overline{\ln p}_{j,t}$ or $\overline{\ln p^{detrend}_{j,t}}$ as dependent variable. Heteroskedasticity-robust standard errors. * indicates significance at the 5% level. Percentage changes in retail prices can be obtained from the following transformation of the estimated parameters: $100(\exp(\beta)-1)$.

where $\overline{\ln p}_{j,t}$ is the average log retail price of product j before and after the alliance formation, β_j are product fixed effects, β_{post} is a fixed effect for the post-alliance period (2015), $\bar{u}_{j,t}$ is an error term, and $t = \{2013, 2015\}$. The parameter β^{in}_{post} measures how the (average) log retail price difference between private labels and national brands sold by a retailer involved in an alliance has changed after the formation of the alliances. Similarly, the parameter β^{out}_{post} measures how the (average) log retail price difference between private labels and national brands sold by either retailer 7 or 8 has changed after the alliances. The OLS estimates of (16) are reported in the first part of Table 10. Results show that the retail prices of national brands sold by a member of an alliance have, on average, decreased by 7.32% relative to that of private labels. This price decrease is, moreover, statistically significant at the 5% level which is reassuring as to the validity of the estimates displayed in Figure 5a.

Interestingly, estimates reported in the first part of Table 10 also indicate that the reaction of retailers 7 and 8 to the alliances is statistically significant with a price decrease of 5.73%, which differs from the estimates displayed in Figure 5b. This contrasting result is explained by the fact that the retail prices of national brands sold by retailers 7 and 8 trend downward in the pre-alliances periods (see Figure 4). By simply averaging the data pre- and post-alliances, the regression model (16) omits to account for such a trend and, in turn, overestimates the price reaction of retailers 7 and 8 to the alliances. To address this issue, I employ a 2-step procedure. First, I detrend the retail prices of national brands sold by retailers 7 and 8 by using the following regression model:

$$\begin{aligned}\ln p_{j,t} = \; & \bar{\beta}\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{pre-alliances}\}_t \\ & + \beta^{trend}_t \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{pre-alliances}\}_t + e_{j,t}\end{aligned} \tag{17}$$

where $\bar{\beta}$ captures the average retail price of national brands sold by retailers 7 and 8 in the pre-alliance periods, β^{trend}_t are fixed effects capturing the downward trend (the last period before the alliances being the base period), and $e_{j,t}$ is an error term. Based on the OLS estimates of (17), I detrend the retail prices of national brands sold by retailers 7 and 8 as follows: $\ln p_{j,t}^{detrend} = \ln p_{j,t} - \hat{\beta}_t^{trend} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} \times \mathbb{1}\{\text{pre-alliances}\}_t$. Then, I average the data pre- and post-alliances and I estimate the regression model (16) using $\overline{\ln p_{j,t}^{detrend}}$ instead of $\overline{\ln p}_{j,t}$ as dependent variable. The OLS estimates reported in the second part of Table 10 indicate a lower decrease in the retail prices of national brands sold by retailers 7 and 8 following the alliance formation (−1.69%). Moreover, the price reaction of retailers 7 and 8 is not statistically significant, which is reassuring as to the validity of the estimates displayed in Figure 5b.

B.2 Alliance-by-alliance event-study analysis

To gain additional insights on the effect of the formation of the three buyer alliances on retail prices, I use an alliance-by-alliance event study. Based on the specification of the regression model (1), I estimate the following reduced-form pricing equation:

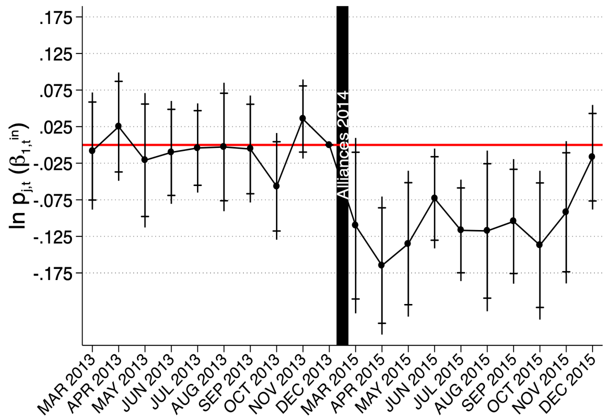
$$\begin{aligned}\ln p_{j,t} = \quad & \beta_j + \beta_t + \beta_{1,t}^{in} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance 1}\}_{j,t} \\ & + \beta_{2,t}^{in} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance 2}\}_{j,t} \\ & + \beta_{3,t}^{in} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance 3}\}_{j,t} \\ & + \beta_t^{out} \mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{retailers 7 \& 8}\}_{j,t} + u_{j,t}\end{aligned} \tag{18}$$

where $\mathbb{1}\{\text{alliance a}\}_{j,t}$, with $a = \{1,2,3\}$, is an indicator variable equals to 1 is product j in market t is sold by a retailer involved in the alliance a. The parameters of interest are the fixed effects $\beta_{1,t}^{in}$, $\beta_{2,t}^{in}$, and $\beta_{3,t}^{in}$ which respectively measure the evolution of the log retail price difference between private labels and national brands sold by retailers involved in the alliances 1, 2, and 3. I estimate the regression model (18) by OLS with standard errors clustered at the product level. Figures 6a, 6b, and 6c display the estimation results.

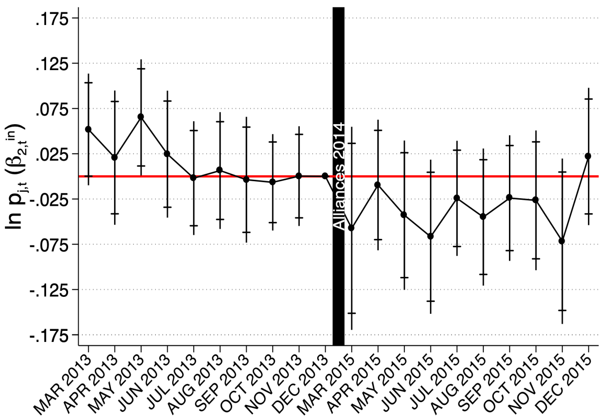
During the pre-alliances periods, every figure shows that the trend in retail prices of national brands sold by retailers involved in each buyer alliance is stable and not statistically different from that of private labels. In the post-alliances periods, Figures 6a and 6c depict a clear trend break where most parameters are negative and significantly different from zero. While the magnitude of the trend break displayed in Figure 6c is similar to that observed in Figure 5a, parameter values shown in Figure 6a are around −0.125. This suggests that the retail prices of national brands sold by retailers 1 and 2 which have formed the alliance 1 have decreased, on average, by 11.75% relative to the retail prices of private labels. Figure 6b also indicates that the retail prices of the retailers involved in the alliance 2 have decreased in the post-alliances periods. However, the estimates suggest that this price decrease is not only of a lower magnitude compared to the other alliances but it is also not statistically significant.

Two main results can be drawn from this alliance-by-alliance event study analysis. First, after the formation of the alliances, the national brands purchased by each buyer alliance have experienced a retail price drop relative to private labels. Second, the magnitude of this price drop varies across alliances. While the retail prices of national brands purchased by the alliances 1 and 3 have substantially declined (around −11.75% and −7.23%, respectively), the retail prices of national brands purchased by the alliance 2 have not significantly decreased relative to the retail prices of private labels.

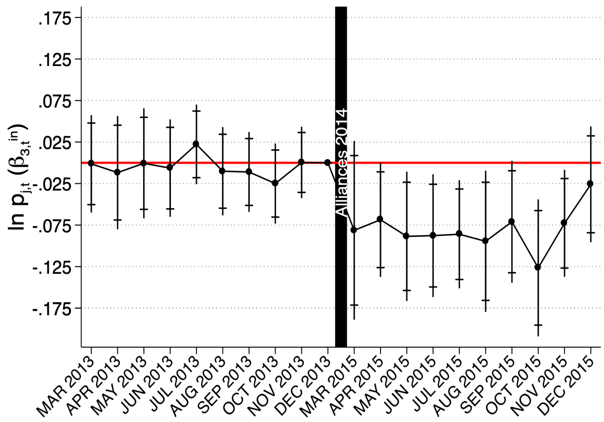
Figure 6: Event study (alliance-by-alliance)



(a) Alliance 1: retailers 1 & 2



(b) Alliance 2: retailers 3 & 4



(c) Alliance 3: retailers 5 & 6

Notes: OLS estimates of the coefficients $\beta^{in}_{1,t}$, $\beta^{in}_{2,t}$, and $\beta^{in}_{3,t}$ from the regression model (18). The number of observations is 2,192 and the R^2 adjusted equals 0.99. Capped-bars and bars indicate respectively the 90% and 95% confidence intervals with standard errors clustered at the product level (111 clusters). The coefficients $\beta^{in}_{1,t}$, $\beta^{in}_{2,t}$, $\beta^{in}_{3,t}$, and β^{out}_{t} for December 2013 are normalized to zero (base period). Percentage changes in retail prices can be obtained from the following transformation of the estimated parameters: $100(\exp(\beta)-1)$.

C Demand results: Additional tables and figures

Table 11: First-stage regression (TSLS)

Variable	Coefficient	S.E.
$\sum_{k\in\mathscr{J}_{r(j),t}\setminus\{j\}} 1$	0.004	0.004
$\mathbb{1}\{\text{national brand}\}_{j,t} \times \mathbb{1}\{\text{alliance}\}_{j,t} \times \mathbb{1}\{\text{post-alliances}\}_t$	-0.025^*	0.004
Brand fixed effects (not shown)		
Retailer fixed effects (not shown)		
Market fixed effects (not shown)		
F-stat	23.35	
F_{eff}	20.73	
R^2 adjusted	0.93	
Number of observations	2,192	

Notes: Heteroskedasticity-robust standard errors. * indicates significance at the 5% level. The variable $\sum_{k\in\mathscr{J}_{r(j),t}\setminus\{j\}} 1$ corresponds to the number of products other than j sold by the same retailer in each market. F_{eff} is the robust F-stat of Montiel Olea and Pflueger (2013). The critical value for testing that the TSLS bias exceeds 5% of the OLS bias is 8.93.

D Counterfactual simulations: Additional tables

Table 12: Results of Scenario A "No Alliance" (by retailers)

	%Δ Retail price	%Δ Margins		%Δ Profit of retailers
		Retailers	Manufacturers	
Inside alliances:				
Retailer 1	−12.02	5.33	−90.95	44.82
	(0.29)	(0.14)	(0.27)	(0.56)
Retailer 2	−15.28	3.18	−92.55	69.26
	(1.10)	(0.11)	(0.29)	(3.68)
Retailer 3	−6.64	0.37	−41.55	6.09
	(0.17)	(0.06)	(0.18)	(0.43)
Retailer 4	−9.48	1.78	−53.52	24.00
	(0.33)	(0.06)	(0.27)	(0.48)
Retailer 5	−7.49	1.88	−74.26	26.11
	(0.38)	(0.11)	(0.21)	(0.68)
Retailer 6	−11.21	2.97	−79.30	48.52
	(0.22)	(0.08)	(0.18)	(0.61)
Outside alliances (national brands only):				
Retailer 7	−3.16	−3.29	−67.03	−11.53
	(0.14)	(0.14)	(0.92)	(0.23)
Retailer 8	−2.01	−0.70	−69.47	−11.18
	(0.15)	(0.04)	(1.18)	(0.33)

Notes: Standard deviations in parenthesis refer to variation across markets. Percentage changes in retail prices of national brands and price-cost margins over national brands relative to the "No Alliance" scenario are calculated using quantity weights. Percentage changes in the profit of retailers over national brands (excluding private labels) relative to the "No Alliance" scenario.

Table 13: Results of Scenario B "Joint delisting decision" (by retailers)

	%Δ Retail price	%Δ Margins		%Δ Profit of retailers
		Retailers	Manufacturers	
Inside alliances:				
Retailer 1	−0.91	−0.84	−5.30	−3.60
	(0.10)	(0.11)	(0.92)	(0.49)
Retailer 2	−14.07	3.55	−87.02	88.49
	(1.45)	(0.26)	(7.56)	(12.07)
Retailer 3	−4.07	1.29	−29.42	12.44
	(0.21)	(0.11)	(2.13)	(1.35)
Retailer 4	−4.72	1.08	−24.57	13.96
	(0.41)	(0.14)	(2.03)	(1.98)
Retailer 5	−3.11	0.75	−27.74	10.29
	(0.19)	(0.11)	(1.89)	(1.31)
Retailer 6	−2.89	0.87	−24.61	12.45
	(0.23)	(0.09)	(1.71)	(1.52)
Outside alliances (national brands only):				
Retailer 7	−1.00	−1.21	−20.62	−4.18
	(0.04)	(0.04)	(0.34)	(0.11)
Retailer 8	−0.63	−0.25	−21.80	−4.12
	(0.05)	(0.02)	(0.41)	(0.15)

Notes: Standard deviations in parenthesis refer to variation across markets. Percentage changes in retail prices of national brands and price-cost margins over national brands relative to the "No Alliance" scenario are calculated using quantity weights. Percentage changes in the profit of retailers over national brands (excluding private labels) relative to the "No Alliance" scenario.

Table 14: Results of Scenario C "Joint-bargaining paradox" (by retailers)

	%Δ Retail price	%Δ Margins		%Δ Profit of retailers
		Retailers	Manufacturers	
Inside alliances:				
Retailer 1	−0.28	−0.31	−1.59	−1.38
	(0.01)	(0.02)	(0.13)	(0.13)
Retailer 2	−3.28	0.74	−20.05	14.35
	(0.22)	(0.04)	(0.80)	(0.63)
Retailer 3	−0.28	−0.18	−1.67	−1.30
	(0.02)	(0.01)	(0.20)	(0.15)
Retailer 4	−3.75	1.32	−21.56	16.09
	(0.12)	(0.04)	(0.31)	(0.25)
Retailer 5	−0.19	−0.21	−1.79	−1.52
	(0.02)	(0.02)	(0.15)	(0.09)
Retailer 6	−3.04	1.03	−21.49	15.16
	(0.06)	(0.04)	(0.24)	(0.23)
Outside alliances (national brands only):				
Retailer 7	−0.31	−0.44	−6.02	−1.55
	(0.01)	(0.02)	(0.27)	(0.07)
Retailer 8	−0.19	−0.09	−6.50	−1.57
	(0.02)	(0.01)	(0.36)	(0.07)

Notes: Standard deviations in parenthesis refer to variation across markets. Percentage changes in retail prices of national brands and price-cost margins over national brands relative to the "No Alliance" scenario are calculated using quantity weights. Percentage changes in the profit of retailers over national brands (excluding private labels) relative to the "No Alliance" scenario.

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