Working Papers SMART

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Working Paper SMART N°22-06

October 2022

UMR **SMART**, INRAE - L'Institut Agro Rennes-Angers

(Structures et Marchés Agricoles, Ressources et Territoires)

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Acknowledgments

We thank Kurt Brekke, Etienne Chantrel, Angela Cheptea, Xavier d'Haultfoeuille, Rachel Griffith, Fabian Herweg, Marc Ivaldi, Patrick Rey and Lars Sorgard, as well as participants to the BECCLE 2019 and 2017 Conferences (Bergen), CRESSE 2019 (Rhodes), EARIE 2019 (Barcelona), ESEM 2017 (Lisbon), Network of Industrial Economists Conference 2022 (Loughborough), Workshop on IO, Trade and the Food Industry 2018 (Paris), Workshop ECODEC on industrial organization 2018 (Paris) and seminars at Bayreuth University, ESCP, Paris School of Economics and University of Nantes for helpful comments. Allain gratefully acknowledges support from a grant of the French National Research Agency (ANR), "Investissements d'Avenir"(LabEx Ecodec/ANR-11-LABX-0047).

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Abstract We develop a model of vertical relations between national brand and private label producers and competing multi-product retailers to derive new predictions on the impact of input price discrimination on retail prices. A reform that lifted a ban on input price discrimination in France provides a natural experiment to test these predictions. Using household scanner data on food prices, we run a difference-in-differences analysis and show that the reform caused a significant decrease of the relative prices of national brand products. These results suggest a pro-competitive effect of authorizing input price discrimination.

Keywords: input price discrimination, policy evaluation, food retail sector.

JEL Classification: K21, L13, L42, L66, L81.

L'effet de la discrimination des prix sur le marché amont sur les prix de détail : théorie et empiriques à partir du cas de la France

Résumé Nous développons un modèle de relations verticales entre les producteurs de marques nationales et de marques de distributeur et des distributeurs multi-produits concurrents pour dériver de nouvelles prédictions sur l'impact de la discrimination des prix sur le marché amont sur les prix de détail. En France, une réforme a levé l'interdiction de la discrimination des prix sur le marché amont et fournit une expérience naturelle pour tester ces prédictions. A partir de données des scanners des ménages sur les prix des denrées alimentaires, nous effectuons une analyse des différences de différence et montrons que la réforme a généré une baisse significative des prix relatifs des produits de marque nationale. Ces résultats suggèrent un effet pro-concurrentiel de l'autorisation de la discrimination des prix sur le marché amont.

Mots-clés: discrimination des prix sur le marché amont, évaluation des politiques, grande distribution alimentaire.

Classification JEL: K21, L13, L42, L66, L81.

1. Introduction

Input price discrimination, also called secondary line discrimination or wholesale price discrimination, characterizes the behavior of a supplier who applies different conditions of sales to buyers who compete to resell to consumers. In theory, input price discrimination arises because it is beneficial for a supplier to exploit downstream firms' heterogeneity in demand or in costs, which results in different price elasticities of demand for the input. As in the case of final price discrimination, the effects of input price discrimination on final prices and welfare are likely to be ambiguous. Furthermore, the vertical relations between suppliers and competing retailers generate additional forces that lead to even more complex effects. In particular, buyers may be tempted to be more aggressive in their negotiations with suppliers when the latter can price-discriminate, as they negotiate for personalized rebates that are not granted to their rivals, which gives them a competitive advantage on the downstream market. This *competition effect* may translate into lower final prices. In contrast, if buyers with a high bargaining power obtain more advantageous input prices than their smaller rivals, the latter may be driven out of the market: this *exclusion effect* may translate into higher final prices. Overall input price discrimination is likely to have ambiguous effects on retail prices and welfare, as confirmed by the literature (see Section 2).

Historically, competition rules regarding input price discrimination were mainly motivated by exclusion concerns. In the US, a seller is prevented from applying dissimilar conditions to equivalent transactions where the effect "may be to lessen competition".[1] The same applies in Europe when the seller is a dominant firm.[2] Going one step further, France adopted in 1986 a specific regulation that forbids any supplier to offer different conditions to similar buyers, before moving backwards in 2008 with the "Loi de Modernisation Economique" (henceforth LME) with the aim of reducing retail prices and increasing the purchasing power of consumers through the *competition effect*. Interestingly, Norway recently considered reaching the same goal by adopting rules diametrically opposed to the French legislation. After an entrant online shopping platform complained that it could not compete on equal terms with incumbent retailers on the food retail market because they were benefiting from lower input prices, Norway considered introducing a ban on input price discrimination in order to limit the *exclusion effect*.[3] The above cases highlight that

[1] The Robinson Patman Act enacted in 1936 prevents "a seller from discriminating in prices among its purchasers for good of like grade and quality" where the effect "may be to lessen competition".

[2] In the EU, Article 102 of the TFEU prohibits a dominant firm from "applying dissimilar conditions to equivalent transactions with other trading parties, thereby placing them at a competitive disadvantage".

[3] See the press release of the Norwegian competition authority at https://konkurransetilsynet.no/investigating-competition-issues-in-the-norwegian-grocery-sector/?lang=en. In 2020, Norway finally decided to impose restrictions, but not a ban on input price discrimination. Note however that, in the EU, there are specific rules regarding discrimination between online and offline retailers. The

the effects of input price discrimination remain an open question for competition policy.

In this paper, we develop both a theoretical and an empirical analysis to investigate the effect of input price discrimination on retail prices. We first design an original model of vertical relations featuring upstream competition between differentiated producers and downstream competition between multi-product retailers. We consider two types of producers: a producer of national brand products who supplies two retailers, and two dedicated private label producers, each of them supplying (exclusively) one of the retailers. As each private label producer has an exclusive relationship with a retailer, the legislation on input price discrimination does not affect their input price. In contrast, we show that a ban may lead to an increase or a decrease of the national brand's input price. The national brand producer may use the ban to commit to a higher input price so as to relax downstream competition: the elimination of the *competition effect* then dominates. The national brand producer may instead use the ban to commit to a low input price to reduce the retailers' *status quo* profit: if he were to sell only his private label, the retailer would then face a more competitive rival. To our knowledge, this *bargaining leverage effect* was not previously highlighted in the literature. Regarding final prices, we show that the price of private labels is likely to be affected as well because the multi-product retailer may divert demand from one product to the other by shifting the two final prices in opposite direction. This effect is henceforth referred to as the *Edgeworth-Salinger effect*. This analysis leads to four scenarios regarding the impact of input price discrimination on final prices: in the first three scenarios the *competition effect* dominates and both the input and final prices of the national brand decrease whereas the final price of private labels either decreases, is unaffected, or even increases due to the *Edgeworth-Salinger effect*; in a fourth scenario, the *bargaining leverage effect* dominates and all prices increase.

This article falls within the theoretical literature on input price discrimination in a secret contracting environment. In line with the seminal article of Hart and Tirole (1990) who highlight that secret contracts trigger opportunism issues, O'Brien and Shaffer (1994) and O'Brien (2014) show that a ban on input price discrimination may solve this issue and restore the monopoly power of a producer supplying competing retailers.[4] In contrast, lifting the ban creates a *competition effect* that leads to a decrease in input prices. In this paper, we consider multi-product retailers – instead of single-product ones – and point out that indirect effects due to multi-product contracting strategies may overturn the results of O'Brien and Shaffer (1994). We show that authorizing price discrimination may instead lead to an increase in the input price prompted by a *bargaining leverage*

draft of the vertical guidelines published in July 2021 by the European Commission explicitly specifies that manufacturers should be able to offer different input prices to online or offline retailers, as long as the differential does not, in practice, foreclose online sales.

[4]See also McAfee and Schwartz (1994).

effect. Moreover, in contrast to O'Brien and Shaffer (1994) who show that final prices go in the same direction as input prices, we point out that the price of some products may go in the opposite direction because of multi-product pricing. This effect indeed relates to the *Edgeworth-Salinger effect* highlighted in the literature on taxation since the seminal paper of Edgeworth (1925).[5] We show that this effect holds in a framework with competing multi-product retailers.

Building on this theoretical approach, we develop an empirical analysis taking advantage of the LME, which lifted the ban on input price discrimination in France, to assess how retail prices are affected by input price discrimination. We contribute to the empirical literature by conducting the first retrospective analysis of the impact of input price discrimination on final prices, based on a quasi-natural experiment. The LME was introduced at the national level and directly applied to all products and retailers. Hence, this quasi-natural experiment does not provide a straightforward control group authorizing a simple estimation of the causal effect of the reform on prices. In light of our theoretical scenarios, our research design relies on the comparison of prices between national brand and private label products over time.

We exploit a rich consumer panel dataset (Kantar Worldpanel, 2010, henceforth KWP), that records all consumer food purchases and prices, at the store level, for a representative set of households in France. We are therefore able to include a large range of food products (approximately 26,000 product-store items), over the 2006-2010 period, which covers two and a half year before and after the adoption of the LME in August 2008. To assess the effect of the LME on retail prices, we compare the mean change in prices of national brand products with that of private label products after the introduction of the LME.

The estimation results show that authorizing input price discrimination caused a decrease of 2.62 percent in the price of national brand products compared to private labels, which comforts the existence of a *competition effect*. This amounts to a decrease by 4.60 euros (out of 175.46 euros) in the monthly average household's shopping basket for national brands. We conduct several robustness tests with different affected groups (*i.e.*, standard private labels, first-price private labels, and discounters' private labels). In all exercises, we find a price decreasing effect of authorizing input price discrimination, ranging from -1.26% to -1.82%. We also provide additional results about the heterogeneity of the price effect of the LME along several dimensions such as time, product categories, and retail groups. We show that the effect materializes right after the passing of the LME, that it affects a large share of food product categories, and that the price decrease is larger at

[5] Edgeworth (1925) highlights that a tax leading to an increase in the price of one of two products sold by a monopolist multi-product retailer may lead to a decrease in the price of the non taxed product, to favor the demand for this product. A similar effect is highlighted by Salinger (1991) who studies the effect of vertical integration – decreasing the input price – on the final prices of a multi-product retailer.

the retailers that were the most expensive before the law.

The paper proceeds as follows. Section 2 reviews the existing theoretical and empirical literature on input price discrimination. Section 3 presents an original model that yields testable predictions about the effect of input price discrimination on prices. Section 4 gives some background on the French food retail sector and the relevant legislation. Section 5 presents the data, discusses the empirical strategy, and derives our main empirical results. Section 6 extends the analysis in several directions, while robustness checks are provided in Section 7, and Section 8 concludes.

2. Related Literature

In this section we present the theoretical and the empirical literature on input price discrimination. The theoretical literature is dense and brings out contrasted results on welfare. In contrast, the empirical literature is scarce but also highlights mixed results. Most of the theory papers consider the case of an upstream monopoly selling one product to competing retailers. The assumptions on contract observability play a key role and we therefore present the main results of the literature according to these assumptions.

☐ Public contracts – In vertically related markets with public linear wholesale contracts, the "standard view" is that banning input price discrimination improves allocative efficiency and welfare (DeGraba, 1990; Katz, 1987). In a setup in which a supplier offers linear contracts to asymmetric competing retailers, the less efficient downstream firm receives a discount because this firm has a more elastic demand for the input. Discrimination thus generates an inefficient allocation of production among producers and harms welfare. Furthermore, DeGraba (1990) shows that this negative effect persists in the long term when firms can invest to lower their marginal cost.

However, the economic literature has developed many arguments against this "standard view". Recent papers have highlighted that these allocative inefficiencies may be reversed when considering more general demand and tariff schemes.[6] For instance, Inderst and Shaffer (2009) show that a supplier offering two-part tariffs charges a lower unit price to the more efficient downstream firm to maximize total industry profit. Under a ban, the monopolist raises both input prices, and relatively more that of the more efficient firm, thus reducing welfare.[7] Katz (1987) and Inderst and Valletti (2009) also show that the

[6] For instance, in a setup with independent downstream markets Gaudin and Lestage (2020) and Li (2017) highlight that the welfare effects of input price discrimination depend on the curvature of the demand function. In a Cournot competition setup, Yoshida (2000) shows that an *increase* in the final good output is a sufficient condition for welfare deterioration, because of production inefficiencies.

[7] Another argument developed by Herweg and Muller (2014) highlights that when retailers are privately informed about their efficiency and that a manufacturer offers a menu of non-linear contracts as a screening

standard view is reversed when downstream firms can integrate backwards as this threat enables the most efficient retailer to obtain the lower wholesale unit price. Focusing on price discrimination across markets rather than across buyers, Miklos-Thal and Shaffer (2019) show that discrimination may have a positive allocative effect, and that welfare can rise.[8]

☐ Secret contracts – Under discrimination with secret contracts and two-part tariffs, fear of opportunism (see Hart and Tirole, 1990) drives each retailer to reject any tariff with a unit input price above the marginal cost. Opportunism thus prevents the monopolistic supplier from capturing the monopoly rent. A ban on discrimination restores the observability of the input price offered to the rival, hence suppressing any scope for opportunism. O'Brien and Shaffer (1994) and O'Brien (2014) show that under such a ban, unit prices increase up to the point where retail prices reach the monopoly level (see also the case of symmetric beliefs in McAfee and Schwartz, 1994). Therefore, under secret contracts, results are less contrasted: input price discrimination triggers a *competition effect* that decreases final prices and increases welfare.[9]

☐ Empirical literature – Although the theoretical literature on input price discrimination highlights mixed results, there are to our knowledge only few empirical papers to cast light on the debate. Among them, Villas-Boas (2009) develops a structural model of demand and supply with public unit wholesale contracts and simulates the effect of banning price discrimination on the wholesale market for coffee in Germany. She finds that imposing uniform input pricing would have welfare improving effects. In contrast, Hastings (2009) models the vertical channel in the gasoline market and simulates equilibrium prices under price discrimination and uniform input pricing. She finds that, on average, final prices would rise five cents per gallon under uniform wholesale pricing. In a secret contracting environment, Grennan (2013) develops a structural model of bargaining and shows that, according to the theoretical predictions, more uniform prices soften competition on the input markets among hospitals.[10]

device, a ban on discrimination has ambiguous welfare effects. Note that such a ban may lead to the *exclusion* of the less efficient retailer.

[8]In the same vein, Arya and Mittendorf (2010) analyze a ban on input price discrimination across asymmetric retailers operating in separate markets, when one of them operates in multiple markets. They find that price discrimination leads to price cuts in markets with lower demand and that, when these markets are also less competitive, price discrimination can provide welfare gains by increasing output on these markets.

[9]An exception is Caprice (2006) who shows that if the monopolist competes with a less efficient competitive fringe, the ban could instead lead to a reduction in the unit price and increase welfare.

[10]Using a similar modeling approach, Yonezawa *et al.* (2020) find evidence of input discrimination in the US yogurt market implying that The Robinson Patman Act is not enforced.

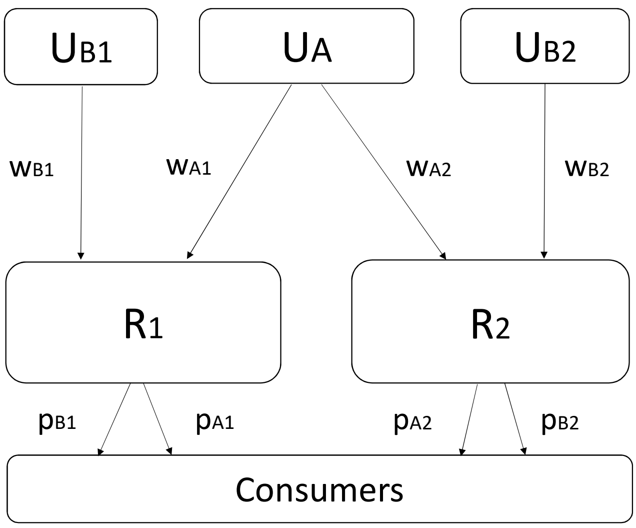
3. The Impact of Input Price Discrimination on Intermediate and Retail Prices

In this section, we build on O'Brien and Shaffer (1994) and O'Brien (2014), and extend their analysis of input price discrimination to take into account upstream competition and multi-product retailers.

3.1. The Model

Consider two differentiated retailers denoted R_i, with $i \in \{1, 2\}$, carrying two differentiated goods $k \in \{A, B\}$. Good A is produced by a manufacturer U_A who sells to the two retailers. Good B is produced by two independent suppliers U_{Bi}. Each supplier U_{Bi} sells its good exclusively to one retailer, R_i (see Figure 1). Good A may represent a national brand, whereas each product Bi may represent the private label product sold by retailer R_i.[11] Overall, there are thus four differentiated products ki available to consumers. For simplicity, we assume that goods A and B are produced at a constant marginal cost $c > 0$.

Figure 1: Market Structure



We assume that consumers' demand for product ki (with $\{k, l\} = \{A, B\}$ and $\{i, j\} = \{1, 2\}$) is twice continuously differentiable in the price vector $(p_{ki}, p_{li}, p_{kj}, p_{lj})$,

[11] It is usual to assume vertical differentiation between private label and national brand products. However, press releases tend to show that the quality of private label products has increased over time, which supports our assumption of horizontal differentiation between private label products and national brands. More precisely, we assume here that the two private label products are homogeneous – hence the differentiation between products $B1$ and $B2$ reflects the retailers' differentiation.

and is symmetric across retailers and across products:

$$D_{ki}(p_{ki}, p_{li}, p_{kj}, p_{lj}) \equiv D(p_{ki}, p_{li}, p_{kj}, p_{lj}).$$

We denote by D_n its derivative with respect to the n_{th} argument. We denote by $D^3_{ki}(p_{ki}, p_{kj}, p_{lj}) \equiv D(p_{ki}, \infty, p_{kj}, p_{lj})$ the demand for product k at retailer R_i when it only sells product k. We make the following assumption.

Assumption 1 *For any given price vector* $(p_{ki}, p_{li}, p_{kj}, p_{lj})$*, demand for product ki is downward sloping, and products are substitutes:*[12]

$$D_1 < 0, D_2 > 0, D_3 > 0$$
$$|D_1| > |D_2|, |D_3| > |D_4|.$$

We consider the following two-stage game:

- Stage 1: Suppliers simultaneously offer to their retailers take-it-or-leave-it secret two-part tariff contracts, consisting of a unit input price w_{ki} and a fixed fee F_{ki}. When discrimination is prohibited, the national brand producer must offer the same unit input price to the two retailers: $w_{A1} = w_{A2} = w_A$. Each retailer then accepts or rejects the offer; a retailer who rejects the offer cannot sell the good.
- Stage 2: The two retailers compete by simultaneously setting their final prices p_{ki}.

Informational structure. We adopt the contract equilibrium concept developed by Crémer and Riordan (1987), which implies that (i) retailers have passive beliefs,[13] and that (ii) U_A sends two independent delegates, who cannot communicate with each other, to make a take-it-or-leave-it offer to each R_i. Retailers do not observe their competitor's input price when they set final prices. When discrimination is prohibited, however, as the national brand producer must offer the same unit input price to the two retailers, this unit price is *de facto* observable by the two retailers.

Discussion. We assume that the ban only affects the variable part of the tariff, namely, w_A. Indeed, in practice, it may be difficult for a court to establish that discriminatory fixed fees have been employed because such fees may compensate personalized services that are difficult to assess, or take the form of rebates or allowances that are by nature opaque and difficult to uncover. Furthermore, this assumption fits well with the French

[12]Note that we do not make any *a priori* assumption on the sign of D_4. Assumption 1 holds when the second argument goes to infinity, hence extends to $D^3_{ki}(p_{ki}, \infty, p_{kj}, p_{lj})$.

[13]When a retailer receives an unexpected offer, he still believes that his competitor received the equilibrium offer (see, *e.g.*, McAfee and Schwartz, 1994; Rey and Tirole, 2007)

case, as the law established a clear distinction between the (non-discriminatory) unit price and the possibly personalized fees. We also discuss further in Section 3.6 the role of suppliers' take-it-or-leave-it offers on our results.

We now determine the sequential equilibrium of this game by proceeding backwards, and assess the impact of a ban on input price discrimination on the equilibrium outcomes.

3.2. Price Competition Stage

Consider first the case in which input price discrimination is allowed. Assume that in stage 1, each R_i accepts contracts (w_{Ai}, F_{Ai}) from U_A and (w_{Bi}, F_{Bi}) from U_{Bi}. R_i's profit is denoted by:

$$\pi^i(p_{Ai}, p_{Bi}, p_{Aj}, p_{Bj}) \equiv \sum_{\{k,l\}\subset\{A,B\}, l\neq k} (p_{ki} - w_{ki})D(p_{ki}, p_{li}, p_{kj}, p_{lj}) - F_{ki} \tag{1}$$

We assume that π^i is twice continuously differentiable, and we denote by π^i_1 the derivative of R_i's profit function π^i wrt. p_{Ai}, π^i_2 its derivative wrt. p_{Bi}, π^i_3 its derivative wrt. p_{Aj} and π^i_4 its derivative wrt. p_{Bj}. The same convention extends to second order derivatives.

We make the following assumptions regarding the profit functions:

Assumption 2 *For a given vector of input prices* $w = (w_{A1}, w_{B1}, w_{A2}, w_{B2})$, *we assume that there exists a unique equilibrium vector of final prices* $p^*(w) = (p^*_{A1}(w), p^*_{B1}(w), p^*_{A2}(w), p^*_{B2}(w))$, *in which* $p^*_{ki}(w)$ *is the equilibrium price of product* ki. *Furthermore, for any vector of positive final prices, and for* $i = 1, 2$,

(i) $0 < \pi^i_{21} \leq -\pi^i_{11}$ *and* $0 < \pi^i_{12} \leq -\pi^i_{22}$.

(ii) $0 \leq -\pi^i_{23} < \pi^i_{13}$ *and* $0 \leq -\pi^i_{14} < \pi^i_{24}$.

(iii) $\pi^i_{31} + \pi^i_{11} < 0$ *and* $\pi^i_{24} + \pi^i_{22} < 0$ *whereas* $\pi^i_{23} + \pi^i_{21} > 0$ *and* $\pi^i_{14} + \pi^i_{12} > 0$.

Assumptions 2 (i) and (ii) are regularity assumptions that ensure in particular the concavity of each retailer's profit function and strategic complementarity in prices.[14] Assumptions 2 (i) to (iii) also ensure that direct price effects always dominate cross price effects on marginal profits.

[14] Part (i) of Assumption 2 ensures that each retailer's profit function is concave in prices, and, with part (ii), that best response functions increase in the rival's prices. These assumptions also ensure that the best response price for each good increases with its own unit cost. See Appendix A.1 for details.

In the competition stage, each R_i maximizes its profit. This yields the following system of first order conditions (henceforth FOC) for each retailer, where the argument is omitted when obvious:

$$\begin{aligned} \pi_1^i &= D_{Ai}(.) + (p_{Ai} - w_{Ai})\frac{\partial D_{Ai}(.)}{\partial p_{Ai}} + (p_{Bi} - w_{Bi})\frac{\partial D_{Bi}(.)}{\partial p_{Ai}} &= 0 \\ \pi_2^i &= D_{Bi}(.) + (p_{Ai} - w_{Ai})\frac{\partial D_{Ai}(.)}{\partial p_{Bi}} + (p_{Bi} - w_{Bi})\frac{\partial D_{Bi}(.)}{\partial p_{Bi}} &= 0 \end{aligned} \tag{2}$$

The first order conditions in system (2) define the best response functions denoted $p_{ki}^r(w_{ki}, w_{li}, p_{kj}, p_{lj})$. Their intersection defines $p^*(w)$. Whether wholesale prices are observable or not, $p^*(w)$ defines the subgame equilibrium final prices because when unobserved, the rival's input prices are consistently anticipated.

Hereafter, we make the following regularity assumption when needed.

Assumption 3 *For any vector of positive prices,*

$$|\pi_{11}^i + \pi_{13}^i| > |\pi_{21}^i + \pi_{23}^i| \quad ; \quad |\pi_{22}^i + \pi_{24}^i| > |\pi_{12}^i + \pi_{14}^i| \quad ; \quad |\pi_{22}^i + \pi_{24}^i| > |\pi_{21}^i + \pi_{23}^i| \quad .$$

This condition ensures that a unit increase in p_{ki} and p_{kj} – due for instance to a cost shock on product k that affects both retailers – affects more a retailer's marginal profit on this product k than his marginal profit on the rival product l. It also ensures that a unit increase in p_{ki} and p_{kj} affects more the retailer's marginal profit on product k than a similar increase in the prices of the other product at both stores (p_{li} and p_{lj}) would.

Note that Assumptions 1 to 3 are satisfied for a wide range of demand functions, and in particular with a linear demand.[15]

Totally differentiating the two FOCs with respect to w_A yields the following proposition:

Proposition 1 *Under Assumptions 1-3, for any vector of wholesale prices that satisfies symmetry across the retailers,* i.e. *such that $w_{A1} = w_{A2} = w_A$ and $w_{B1} = w_{B2} = w_B$, we have $\frac{dp_{Ai}^*}{dw_A} > 0$, whereas the sign of $\frac{dp_{Bi}^*}{dw_A}$ is ambiguous. In the linear case, $\frac{dp_{Ai}^*}{dw_A} > 0$ and $\frac{dp_{Bi}^*}{dw_A} = 0$.*

Proof. See the Appendix A.3. ■

[15] In Appendix A.2, we illustrate our results for the following linear demand specification, in which the parameter $b \in [0,1]$ measures retail substitution (a proxy for intra-brand competition), while the parameter $a \in [0,1]$ reflects product substitution (a proxy for inter-brand competition):

$$D(p_{ki}, p_{li}, p_{kj}, p_{lj}) = \frac{1 - p_{ki} - b(1 - p_{kj}) - a(1 - p_{li} - b + bp_{lj})}{(1 - a^2)(1 - b^2)}.$$

Proposition 1 results from the following trade-off. On the one hand, a decrease in the input price of product A, w_A, drives the retail price of that product down; hence by strategic complementarity the price of product B may decrease as well. On the other hand, the retailer also has an incentive to divert the demand from product B toward product A on which it makes a relatively higher margin: this drives the price of product B up and by retro-action the price of A too.

This result relates to the *Edgeworth-Salinger effect.* In the seminal paper by Salinger (1991), in reaction to the decrease of one input price, in equilibrium a monopolist multi-product retailer can either decrease the prices of both goods, increase the prices of both goods, or decrease the price of the product whose input price decreased, and increase that of the rival.[16] By showing that the effect of w_A on p_{Bi} is ambiguous, our Proposition 1 thus (partly) extends the result of Salinger (1991) to a setting with competing multi-product retailers. We find that the effect of w_A on p_{Ai} is unambiguously positive: this derives from our assumption of demand symmetry.[17]

Lemma 1 *For any vector of wholesale prices that satisfies symmetry across the retailers,* i.e. *such that* $w_{k1} = w_{k2} = w_k$*,*

(i) whenever $\frac{dp^*_{Bi}}{dw_A} \geq 0$*, we have* $|\frac{dp^*_{Bi}}{dw_A}| < |\frac{dp^*_{Ai}}{dw_A}|$*;*

(ii) whenever $\frac{dp^*_{Bi}}{dw_A} < 0$*, for any vector of wholesale prices that also satisfies symmetry across products,* i.e. *such that* $w_A = w_B$*, we have* $|\frac{dp^*_{Bi}}{dw_A}| < |\frac{dp^*_{Ai}}{dw_A}|$.

Proof. See the Appendix A.4. ■

When wholesale prices satisfy symmetry across the retailers, indirect effects of w_A on p_{Bi} are thus likely to be smaller than direct effects on p_{Ai}. Note that by continuity, when the input prices of the two goods are close enough, direct effects are larger than indirect effects.

3.3. Contract Stage under Discrimination

We now determine the contract equilibrium when discrimination is allowed. Let w^d denote the anticipated equilibrium vector of wholesale prices under discrimination $w^d = (w^d_{A1}, w^d_{B1}, w^d_{A2}, w^d_{B2})$.

[16] Note that Edgeworth (1925) studies the effect of an increase in input price caused by a tax, while Salinger (1991) studies the effect of a decrease in the input price following the vertical integration between the retailer and one supplier.

[17] This is in line with Salinger (1991), who mentions that, in case of symmetry, the prices of both goods cannot increase. Luco and Marshall (2020) find empirical evidence of the existence of such an Edgeworth-Salinger effect in vertical merger cases in the US carbonated-beverage industry.

Consider first the offer made by supplier U_{Bi} to retailer R_i. U_{Bi} and R_i anticipate that R_j sets the equilibrium prices $p^*_{Bj}(w^d)$ and $p^*_{Aj}(w^d)$ defined by equation (2), hence the continuation equilibrium in which R_i sets the prices $p^r_{Ai}(w^d_{Ai}, w_{Bi}, p^*_{Aj}(w^d), p^*_{Bj}(w^d))$ and $p^r_{Bi}(w_{Bi}, w^d_{Ai}, p^*_{Aj}(w^d), p^*_{Bj}(w^d))$. U_{Bi} thus maximizes the following profit:

$$\begin{aligned} &\underset{w_{Bi},F_{Bi}}{Max}\,(w_{Bi}-c)D_{Bi}(p^r_{Bi}, p^r_{Ai}, p^*_{Bj}(w^d), p^*_{Aj}(w^d)) + F_{Bi} \\ &s.t. F_{Bi} \leq (p^r_{Ai} - w^d_{Ai})D_{Ai}(p^r_{Ai}, p^r_{Bi}, p^*_{Aj}(w^d), p^*_{Bj}(w^d)) \\ &+ (p^r_{Bi} - w_{Bi})D_{Bi}(p^r_{Bi}, p^r_{Ai}, p^*_{Bj}(w^d), p^*_{Aj}(w^d)) - F_{Ai} - \overline{\pi}^i_A. \end{aligned} \tag{3}$$

In the above equation, $\overline{\pi}^i_A$ denotes the *status quo* profit of R_i when he rejects U_{Bi}'s offer and accepts the equilibrium offer by U_A – in this continuation equilibrium, R_i only sells product A at a price $p^3_{Ai} \equiv p^r_{Ai}(w^d_{Ai}, +\infty, p^*_{Aj}(w^d), p^*_{Bj}(w^d))$, and his profit $\overline{\pi}^i_A \equiv (p^3_{Ai} - w^d_{Ai})D^3_{Ai}(p^3_{Ai}, p^*_{Aj}(w^d), p^*_{Bj}(w^d)) - F_{Ai}$ is independent of w_{Bi}.

In equilibrium, the participation constraint of R_i is binding. Therefore U_{Bi} maximizes his joint profit with retailer R_i with respect to w_{Bi}.

Consider now the offer made by U_A to R_i. Recall that U_A is assumed to send distinct delegates to the two retailers, U_A thus maximizes the following profit (arguments are omitted when obvious):

$$\begin{aligned} &\underset{w_{Ai},F_{Ai}}{Max}\,(w_{Ai}-c)D_{Ai}(p^r_{Ai}, p^r_{Bi}, p^*_{Aj}(w^d), p^*_{Bj}(w^d)) + F_{Ai} \\ &\quad +(w^d_{Aj} - c)D_{Aj}(p^*_{Aj}(w^d), p^*_{Bj}(w^d), p^r_{Ai}, p^r_{Bi}) + F^d_{Aj} \\ &s.t. F_{Ai} \leq (p^r_{Ai} - w_{Ai})D_{Ai}(.) + (p^r_{Bi} - w^d_{Bi})D_{Bi}(.) - F_{Bi} - \overline{\pi}^i_B, \end{aligned} \tag{4}$$

where the *status quo* profit of R_i when he rejects U_A's offer $\overline{\pi}^i_B$ is independent on w_{Ai}. In equilibrium, each retailer's participation constraint is binding: $F_{Ai} = (p^r_{Ai} - w_{Ai})D_{Ai}(.) + (p^r_{Bi} - w^d_{Bi})D_{Bi}(.) - F_{Bi} - \overline{\pi}^i_B$. Again, supplier U_A sets each unit price w_{Ai} to maximize its joint profit with R_i, which boils down to the following program (omitting the argument when obvious):

$$\underset{w_{Ai}}{Max}(p^r_{Ai} - c)D_{Ai}(.) + (p^r_{Bi} - w^d_{Bi})D_{Bi}(.). \tag{5}$$

We thus obtain the following proposition:[18]

Proposition 2 *When discrimination is allowed, under Assumptions 1-3, there is a unique symmetric equilibrium, in which wholesale prices are cost-based:* $w^d_{B1} = w^d_{B2} = w^d_{A1} = w^d_{A2} = c$.

Proof. See Appendix A.5. ■

[18] This result is standard in the literature on interlocking relationships with two-part tariffs. See for instance Allain and Chambolle (2011), Rey and Vergé (2019).

The following corollary derives from proposition 2 and Assumption 2:

Corollary 1 *When input price discrimination is allowed, under Assumptions 1-3, there is a unique symmetric equilibrium in which the retail price is* $p^* \equiv p^*_{ki}(c,c,c,c)$ *for* $k = A, B$ *and* $i = 1, 2$.

3.4. Contract Stage under a Ban on Discrimination

Consider the offer made by supplier U_{Bi} to retailer R_i. We denote by w_A^{nd} the non-discriminatory wholesale unit price offered by U_A, and the anticipated equilibrium wholesale price vector by $w^{nd} = (w_A^{nd}, w_{Bi}^{nd}, w_A^{nd}, w_{Bj}^{nd})$. The pair $U_{Bi} - R_i$ now anticipates that R_j sets the equilibrium prices $p^*_{Bj}(w^{nd})$ and $p^*_{Aj}(w^{nd})$ and that R_i adapts its prices according to $p^r_{Bi}(w_{Bi}, w_A^{nd}, p^*_{Bj}(w^{nd}), p^*_{Aj}(w^{nd}))$ and $p^r_{Ai}(w_A^{nd}, w_{Bi}, p^*_{Aj}(w^{nd}), p^*_{Bj}(w^{nd}))$. Supplier U_{Bi}'s program is the following:

$$\begin{aligned} &\underset{w_{Bi},F_{Bi}}{Max}\,(w_{Bi}-c)D_{Bi}(p^r_{Bi},p^r_{Ai},p^*_{Bj}(w^{nd}),p^*_{Aj}(w^{nd}))+F_{Bi} \\ &s.t. F_{Bi} = (p^r_{Ai}-w_A^{nd})D_{Ai}(.)+(p^r_{Bi}-w_{Bi})D_{Bi}(.)-\overline{\pi}^i_A - F_{Ai}, \end{aligned} \tag{6}$$

where the arguments are omitted when obvious. The outside option profit $\overline{\pi}^i_A$ does not depend on w_{Bi}; each U_{Bi}'s program thus remains similar to equation (3) under the ban on discrimination, and this leads to the following proposition:

Proposition 3 *When discrimination is banned, under Assumptions 1-3, there is a symmetric equilibrium, in which* $w^{nd}_{Bi} = c$ *for* $i = 1, 2$.

Proof. Supplier U_{Bi}'s program is not affected by the ban on discrimination. ∎

Consider now the contracts offered by U_A to the two retailers. Under the ban, when R_i receives a contract, he knows that his rival receives a contract with the same wholesale unit price. The wholesale unit price in the contract is thus no longer secret. Moreover, from Proposition 3, we have $w^{nd}_{B1} = w^{nd}_{B2} = c$ and this is anticipated by all. For $(w^{nd}_{Bi}, w^{nd}_{Bj}) = (c, c)$, U_A and each R_i thus anticipate the continuation equilibrium in which both R_1 and R_2 adapt their prices to w_A, *i.e.* set $p^*_{A2}(w_A) = p^*_{A1}(w_A) \equiv p^*_{A1}(w_A, c, w_A, c)$ and $p^*_{B2}(w_A) = p^*_{B1}(w_A) = p^*_{B1}(w_A, c, w_A, c)$.

U_A's program with R_1 (and symmetrically for R_2) is the following:

$$\begin{aligned} &\underset{w_A,F_{A1}}{Max}\,(w_A-c)D_{A1}(p^*_{A1}(w_A),p^*_{B1}(w_A),p^*_{A2}(w_A),p^*_{B2}(w_A))+F_{A1} \\ &+(w_A-c)D_{A2}(p^*_{A2}(w_A),p^*_{B2}(w_A),p^*_{A1}(w_A),p^*_{B1}(w_A))+F_{A2} \\ &s.t. F_{A1} \leq (p^*_{A1}(w_A)-w_A)D_{A1}(.)+(p^*_{B1}(w_A)-c)D_{B1}(.)-\overline{\pi}^1_B(w_A). \end{aligned} \tag{7}$$

The *status quo* profit $\overline{\pi}_B^1(w_A)$ now depends on w_A. Indeed, when R_1 does not distribute product A, he knows that R_2 still sells product A facing the same w_A; hence the demand for product B at retailer R_1 may depend on w_A. Furthermore, the participation constraints of the retailers are binding; hence U_A's program simplifies as follows:

$$\underset{w_A}{Max} \sum_{i=1,2}(p^*_{Ai}(w_A)-c)D_{Ai}(.)+\sum_{i=1,2}(p^*_{Bi}(w_A)-c)D_{Bi}(.)-\sum_{i=1,2}\overline{\pi}^i_B(w_A). \tag{8}$$

We rearrange the associated FOC and compute its value in $w_A = w_{Bi} = c$ and in the corresponding equilibrium price p^*. We show that the equilibrium wholesale price w_A^{nd} is higher than c if and only if the sign of the following expression is positive:[19]

$$\tfrac{dp^*_{Ai}}{dw_A}\underbrace{[\pi^i_3-\bar{\pi}^i_3]}_{\phi}+\tfrac{dp^*_{Bi}}{dw_A}\underbrace{[\pi^i_4-\bar{\pi}^i_4]}_{\psi}. \tag{9}$$

The sign of equation (9) hinges on two terms, ϕ and ψ.

First, $\phi \geq 0$ means that the marginal benefit of an increase in the price of product A at retailer j through the diversion of demand is larger for retailer i when he sells both products rather than only product B.[20] Hereafter, we make the following assumption.

Assumption 4 *When $w_A = w_B = c$, at the continuation equilibrium final prices, $\phi \equiv \pi^i_3 - \bar{\pi}^i_3 > 0$.*

Assumption 4 is likely to hold for a wide range of demand functions because an increase in p_{Aj} affects more the demand for the same product than the demand for the other product at the rival retailer i. In particular, it is satisfied with our linear demand specification (see Appendix A.6).

In contrast, ψ compares the marginal benefit of an increase in the price of product B at the rival retailer j for retailer i, when he sells both products, and when he sells only product B.[21] The sign of ψ is ambiguous.[22] In Appendix A.6, we thus develop the analysis without restriction regarding the sign of ψ.

We obtain the following proposition:

[19] See Appendix A.6 for more details.

[20] Formally $\phi = (p^* - c)(\frac{\partial D^*_{Ai}}{\partial p_{Aj}} + \frac{\partial D^*_{Bi}}{\partial p_{Aj}}) - (p^3_{Bi} - c)\frac{\partial D^3_{Bi}}{\partial p_{Aj}}$.

[21] Formally $\psi = (p^* - c)(\frac{\partial D^*_{Ai}}{\partial p_{Bj}} + \frac{\partial D^*_{Bi}}{\partial p_{Bj}}) - (p^3_{Bi} - c)\frac{\partial D^3_{Bi}}{\partial p_{Bj}}$.

[22] Note that with our linear demand specification, ψ is negative, but $\frac{dp^*_{Bi}}{dw_A} = 0$, while $\frac{dp^*_{Ai}}{dw_A} = \frac{1}{2-b} > 0$; hence the sign of equation 9 is that of Φ, which is positive.

Proposition 4 *Under Assumptions 1-4, a ban on discrimination leads to $w_A^{nd} > c$, except when $\psi < 0$ and $\frac{dp^*_{Bi}}{dw_A} > 0$ – in the latter case, the effect of the ban on w_A^{nd} is ambiguous. In the linear specification, we have $w_A^{nd} > c$.*

Proof. See Appendix A.6. ■

The ban on discrimination enables the supplier to solve the opportunism problem and improve the joint profit. Increasing w_A thus benefits the supplier through this *competition effect*. However, a *bargaining leverage* effect is also at play, as an increase in w_A also improves the retailer's *status quo* profit. Decreasing w_A thus reduces the retailer's *status quo* profit. However, this bargaining leverage effect can only occur if (i) $\frac{dp^*_{Bi}}{dw_A} > 0$ and (ii) $\psi < 0$. If at play, the *bargaining leverage effect* can countervail the *competition effect* on joint profit and dampen U_A's incentive to increase its input price. Should the *bargaining leverage effect* dominate the *competition effect*, the equilibrium w_A would be set below c. However, as explained in Appendix A.6, conditions such that equation (9) is negative are quite difficult to meet.

Corollary 2 *All input prices being anticipated, under Assumptions 1-3, there exists a retail price equilibrium symmetric across retailers but asymmetric across products, with $\hat{p}_A \equiv p^*_{Ai}(w_A^{nd}, c, w_A^{nd}, c)$ and $\hat{p}_B \equiv p^*_{Bi}(w_A^{nd}, c, w_A^{nd}, c)$.*

3.5. The Effect of Authorizing Input Price Discrimination on Retail Prices

Comparing propositions 3 and 4, we obtain that authorizing input price discrimination has no effect on the input price of private label products, $w_B^{nd} = c$; however it leads to a change in the unit wholesale price of the national brand product. Note first that, due to our symmetry assumption (see the discussion of proposition 1), in equilibrium, p_A always varies in the same direction as w_A. Table 1 summarizes the predictions regarding the effect of authorizing input price discrimination on final prices that derive from our theoretical findings (propositions 1 to 4). It highlights four possible scenarios.

Table 1: Potential Effects of Authorizing Input Price Discrimination

	$w_A \nearrow$ *Bargaining leverage effect*	$w_A \searrow$ *Competition effect*
$\frac{dp^*_{Bi}}{dw_A} = 0$	–	(i) $p_A \searrow$ $p_B \rightarrow$
$\frac{dp^*_{Bi}}{dw_A} < 0$	–	(ii) $p_A \searrow$ $p_B \searrow$ (iii) $p_A \searrow$ $p_B \nearrow$ *Edgeworth-Salinger effect*
$\frac{dp^*_{Bi}}{dw_A} > 0$	(iv) $p_A \nearrow$ $p_B \nearrow$	(ii) $p_A \searrow$ $p_B \searrow$

In the first three scenarios, the *competition effect* dominates the *bargaining leverage effect*, and authorizing input price discrimination results in a decrease in w_A and p_A:

- Scenario (i): whenever $\frac{dp^*_{Bi}}{dw_A} = 0$, the input and final prices of private labels are unaffected. This scenario occurs for instance in the linear demand specification. It also coincides with the theoretical predictions of O'Brien and Shaffer (1994) in the case of a monopolist supplier.

Our theoretical analysis however highlights that the price of private labels may be indirectly affected ($\frac{\partial p^*_{Bi}}{\partial w_A} > 0$ or $\frac{\partial p^*_{Bi}}{\partial w_A} < 0$);

- Scenario (ii): the prices of both national brands and private labels decrease. The *competition effect* extends from national brands to private labels.

- Scenario (iii): the price of national brands decreases, but that of private labels increases. Due to the *Edgeworth- Salinger effect*, the retailer increases its margin on private label product to divert demand towards national brands.

Finally, whenever the *bargaining leverage effect* dominates the *competition effect*, w_A and thus p_A increase, leading to the following scenario:

- Scenario (iv): all final prices increase. Note that we know from Lemma 1 that in such a case, the final prices of private label products increase less than the final prices of national brand products.

In terms of policy evaluation, these scenarios lead to different conclusions. Under scenarios (i) and (ii), a lift of the ban on input price discrimination, such as the the LME reform, would reach its target and unambiguously increase consumer surplus. Under Scenario (iii), the effect would be less clear as the increase in private label prices may countervail the decrease in national brand prices, hence leading to a potential ambiguous effect on consumer surplus. An important consequence of Lemma 1 is that, under Assumptions 1 to 4, when wholesale price variations are small enough following the lift of the ban, p_B varies less than p_A in absolute terms. This result implies that scenario (iii) is more likely to lead to an increase in consumer surplus, given that private labels also represent a lower share of consumers expenditures than national brands. Under the last scenario (iv), all prices would increase, and the reform would miss its target and actually harm consumer surplus.

Overall, our model predicts a differentiated impact of authorizing input price discrimination on the final prices of the national brand and private label products. We build our empirical strategy (developed in Section 5) on this prediction, conducting a difference-in-differences analysis on final prices and taking private labels as a comparison group.

3.6. Robustness and Discussion

We first discuss our assumption that the suppliers of product B offer take-it-or-leave-it contracts to retailers in stage 1. We then explore how our results would change if retailers had some bargaining power.

3.6.1. Production of the Private Label

In the literature on private labels, suppliers of private labels are often considered as vertically integrated with a retailer (retailers operate their own manufacturing plants), or produced by a competitive fringe of independent, small manufacturers.[23] Assuming in-house production by the retailer, or production by a competitive fringe, would also lead to cost-based tariffs whether input price discrimination is allowed or not. Alternatively, private labels may also be produced by large manufacturers who produce both their national brand and private labels. Note that even if a producer (a national brand producer) was producing private labels for several retailers, the ban would not affect their input prices, as each private label is designed for a specific retailer and thus considered as a different product. Furthermore, under our contract equilibrium assumption, if the same producer were supplying the national brand and the private label product to competing retailers, the ban would not affect the input pricing strategy of this supplier for the private label.

3.6.2. Bargaining Assumptions

We discuss here how the equilibrium outcomes are affected by the contracting setup, and especially how they would be affected by the introduction of bargaining between suppliers and retailers. Note that it is difficult to model bargaining in a context in which suppliers are submitted to the ban on price discrimination, which somehow implies that retailers are price-takers.

One possibility to circumvent this issue is to follow O'Brien and Shaffer (1994), who consider that suppliers still offer take-it-or-leave-it unit prices whereas a bargaining takes place over a fixed fee. As the ban we study does not prevent discrimination on the fixed fees, this setup would lead to the same predictions as our baseline model.

Another possibility is to instead assume as O'Brien (2014) that the supplier publicly enters in a bargaining with one retailer and then applies the resulting contract to its rival. In case of a breakdown in the negotiation between U_A and, say, D_1, then the supplier would renegotiate with D_2 as a monopolist and agree on different contract terms. In

[23] See Chambolle *et al.* (2015) for a discussion on the organizational choice of production of private labels.

such a setup, the input price has no effect on the *status quo* profit of the firms (which derives from the renegotiation), and thus the ban unambiguously leads to an increase in the wholesale unit price regardless of the distribution of bargaining power. Therefore in this set-up, scenario (iv) would be eliminated.

4. A Natural Experiment: The French Grocery Sector

Over the last decades, the regulatory environment of the contractual relationships between manufacturers and retailers has undergone several changes in France (see below). We exploit the (quasi-)natural experiment offered by the LME (enacted in 2008) to evaluate how authorizing input price discrimination affects final prices. We leverage a rich dataset on food purchases that covers two and a half years before and after the introduction of the LME, spanning from 2006 to 2010. This section briefly describes the main features of the French food retail sector (Section 4.1), before presenting the legal context and the changes implemented by the LME (Section 4.2).

4.1. The French Grocery Sector

In 2008, the French grocery sector represented about 72.5% of total food purchases (source: INSEE, 2011) with a retail network of about 15,000 stores. It was highly concentrated, with a cumulative market share of 87% for the six largest groups at the national level, and a much higher concentration at the local level.[24] Although each retail group operates several retail chains, negotiations take place at the national level between the retail groups (or often alliances of retail groups) and their suppliers. On the suppliers' side, a few large groups, such as Unilever, PepsiCo, and Danone, represent a large share of the total added value in the food chain, but 98% of suppliers in the food industries are small and medium-sized enterprises. Therefore, the balance of power between manufacturers and retailers is often in favor of the buyers, and the press regularly reports sharp tensions during the annual negotiations, which take place from November to February.[25] These negotiations determine the tariff, which consists in a wholesale unit price, and usually includes several types of rebates and fees, such as slotting fees, or fees for services undertaken by retailers (*e.g.*, promotional operations, market studies,...). In these negotiations, an important source of buyer power for retailers is the growing share of their private labels. In France, the average market share of private labels for food products was about 30% over the period of study (see Appendix B.2).

[24]The market shares are computed using KWP data. See (Autorité de la concurrence, 2010, par.29), and Allain *et al.* (2017) for detailed information on the French local market structure.

[25]Products negotiated on spot markets, such as fresh fruits and vegetables, meat, or fish, are not concerned by these annual negotiations.

4.2. The Legal Framework of Negotiations

Since 1986, price discrimination by a supplier between similar buyers was forbidden in France. In particular, producers had to publish "general terms of sales" that had to be identical for all their "similar" buyers. In practice, this ban on input price discrimination ensured that two retailers obtained the same unit wholesale price from a supplier, but they could still pay different fixed fees, which were quite difficult to evaluate. The opacity surrounding the negotiations made it difficult for a court to establish whether discriminatory fixed fees had been used.

More than twenty years later, the *Loi de Modernisation Economique* (LME) revoked this ban and authorized suppliers to price-discriminate between retailers. This reform, passed on August 5, 2008, was part of a global attempt to intensify competition among retailers in order to increase the purchasing power of consumers. In particular, another process of reforms had already started in 2005 to break-up the inflationary mechanism caused by the former legislation, the 1996 Galland Act. This Act prevented retailers from setting retail prices below a threshold defined as the unit price invoiced by the supplier. This threshold excluded all conditional and deferred rebates. Moreover, as price discrimination was not allowed, this threshold was common to all retailers. Allain and Chambolle (2011) show how the conjunction of banning both resale below cost and price-discrimination actually turned the price threshold into a uniform price-floor that neutralized retail competition, leading to higher final prices.[26] In 2005, the Dutreil Act started breaking-up this mechanism by enabling retailers to incorporate most of these rebates in the resale below-cost threshold.[27] The Châtel Act (January 2008) finalized the reform by allowing retailers to include all types of rebates in the threshold. Although this reform took place only a few months before the LME reform, we are confident that most of the effect of the Galland Act process came with the Dutreil Act of 2005.[28]

5. Empirical Strategy and Main Results

The central prediction of our model is that authorizing input price discrimination should impact differently the final prices of national brands and private labels. We assess the price effect of the LME by comparing the mean change in prices of national brand products (henceforth NB products) with that of private label products (henceforth PL products)

[26]Biscourp et al. (2013) highlight that the correlation between local market concentration and retail prices collapsed after the Galland Act, which they interpret as final prices becoming more uniform across markets.

[27]Specifically, all rebates representing more than 15% of the invoiced unit price were integrated in the new threshold.

[28]See for instance Allain *et al.* (2008).

after the introduction of the LME. We choose August 2008 as the first month under the new regulation as input price discrimination became lawful immediately after the passing of the law, and this led some retailers to renegotiate the terms of their 2008 contracts without delay.[29]

This research design is similar to a difference-in-differences analysis (henceforth DiD analysis). However, the key assumption of a DiD analysis, that the control group is not affected by the event under study may not be satisfied in our case. Indeed, we have shown in the theory section that retailers may have reacted to the lift of the ban by adjusting (up or down) the price of private labels. This is the case in scenario (ii), (iii) and (iv) in Section 3.5. In such cases, the estimated coefficient gives a biased measure of the causal effect of the law on the price of the NB products, hence on prices in general. We show further how our estimation results enable us to discriminate among the different price scenarios delivered by the model, which, in turn, help us understand whether and in which direction our estimates are likely to be biased.

Section 5.1 presents the data and how we proceed to construct the variables of interest and the final sample. In Section 5.2, we report some preliminary results on the price evolution of the groups of products under study. Section 5.3 presents the results of the DiD analysis and assesses the relative change in prices of NB products resulting from the lift of the ban. Finally, Section 5.4 discusses the interpretation of the empirical results in light of the model predictions.

5.1. Data and Sample Selection

5.1.1. Household Scanner Data

We exploit a rich dataset of household food purchases to analyze the price evolution of a wide range of products over the 2006-2010 period. The KWP dataset records information on daily food purchases over a panel of, on average, more than 10,000 households representative of the French population. Purchase data are collected by the households themselves, usually by means of a home scanner.[30] The households record information on the quantity and the expenditure for each purchased product, as well as the store type (*e.g.*, supermarket, discounter, specialized store), and the retail chain where the purchase was made. Furthermore, for products with a European Article Number (EAN, a 13-digit barcode that is a superset of the Universal Product Code), the dataset contains detailed

[29] Several press articles reported these renegotiations. See for instance the interview of the CEO of Leclerc, one of the leaders of the market, in *Le Parisien*, 09/01/2008.

[30] The households can also record their purchases online or through a palm PDA device.

information on product characteristics, including the brand name and the name of the manufacturer – hence a product is described by a unique identifier. Finally, products are clustered into 349 categories of food products, which are aggregated into 61 families of products.

Overall, the KWP data thus offers a unique opportunity to track over time prices and quantities for a wide variety of precisely defined products. We provide additional information on the KWP data in Appendix B.1.

5.1.2. Variables of Interest

Our identification strategy requires information about the brand type of each purchased product. We adopt a fine categorization and classify each brand from the KWP data in 4 brand types following the usual denomination in France: national brand (NB), private label (PL), discount private label (PL-D), and first-price brand (FP). While national brands and private labels (PL and PL-D) correspond to the two upper tiers of the price range, first-price products refer to the lowest tier of a category.[31] A detailed presentation of the brand type classification method is provided in Appendix B.2.

Using the purchase transactions, we compute a mean unit price for each retail chain-product pair on the French market.[32] To ensure a sufficient number of purchase observations per retail chain-product pair, we aggregate the data over 4 weeks.[33] Precisely, the average unit price of a product purchased in a given chain in France during a month is calculated as the ratio of total sales to total quantities. The average unit prices are then expressed in euro per measurement unit (*i.e.*, per kg, per liter or per unit), and are deflated by the monthly consumer price index to abstract from the global trend on prices in the economy.

5.1.3. Sample Selection

We construct the final sample by applying six selection criteria to the food purchase transactions recorded in the KWP. These selection criteria are discussed below.

[31] A first-price product can be either sold under a retailer's brand in all the stores of a given retail group (denoted PL-FP, hereafter) or produced by an independent manufacturer and sold under a generic name in different retail groups.

[32] Data are trimmed to exclude transactions which have a standardized deviation of the unit price from the monthly average price greater than 10 in absolute value. The results are robust to using other cutoffs (see the Online Appendix).

[33] In the terminology of KWP, a month is a 4-weeks period, and a year is thus composed of 13 months.

The Retail Chain Selection. In order to limit some confounding effects resulting from sample composition, we exclude from the sample the distribution channels that are not offering private labels for food products.[34] In short, the sample only contains food purchases made in supermarket chains and their associated online food platforms. A list of the retail groups selected in the final sample and their respective market shares are given in the Online Appendix.

Product Category under the Scope of the LME. We remove from the dataset product categories that are not concerned by the LME because their prices are not set through bilateral contractual relationships but depend on other sales agreements or market mechanisms. Applying this criterion mainly leads us to remove raw agricultural products (*e.g.* fresh meat, fish, fruits and vegetables) traded on spot markets.

National Brands and Private Labels. We keep in our sample all NB products that are sold by at least two retail groups.[35] In contrast, we only keep private labels sold by a single retail group.[36]

Chain-Product Pair over Time. We impose that each chain-product pair must be present at least once in the pre- and once in the post-LME periods. We thus retain 32% of the chain-product pairs identified in the dataset, which represent almost 83% of the observations (defined at the chain-product-month level) these products are consistently offered.[37] From an econometric point of view, this criterion allows us to introduce chain-product fixed effects in the regression model, which control for a potential omitted variable bias related to time-invariant chain-product characteristics.[38]

Product Category Assignment. We require that each product category retained in the final sample is composed of national brands and private labels. As before, the rationale for this criterion is to limit some endogeneity bias resulting from a correlation between

[34]We thus remove from the dataset all product purchases made in non-food distribution channels (*e.g.*, gasoline stations), in specialized distribution channels (*e.g.*, farmer markets, frozen retail chains), as well as in specialized shops (*e.g.*, butchers, bakers, wine merchants), because these retailers do not sell private labels.

[35]We exclude national brands that are offered by a single retail group because those may not be affected by the law. As mentioned in Section 4.1, retailers bargain with suppliers at the group level and not at the chain level.

[36]We thus remove the seldom occurrences of private labels offered in two competing retail groups.

[37]Interestingly, 50% of chain-product pairs are offered for less than a year.

[38]It could be argued that products are not randomly affected by the law, and then that assignment to a group is confounded with the price variable. If product characteristics that affect prices vary significantly across the affected and comparison groups, *e.g* if the group of national brands is mainly composed of organic products whereas private labels are mainly low-quality products, the estimate of the relative change in prices will be biased. Introducing product-chain fixed effect allows to control for this bias.

the passing of the law and some specific product category shocks that may affect retail prices (*i.e.*, some confounding effects at the category level).

Parallel Trends Assumption. The central identification assumption of our DiD analysis is that absent the law, the prices of national brands and private labels would have evolved identically. This assumption cannot be verified empirically over the whole period of study, but it is essential to ensure that it is satisfied in the pre-LME period. We proceed in two steps.

First, our results may be biased if some chains - especially discounters, who sell mostly PL products - were exposed to specific demand and supply shocks, that would impact differently the affected and comparison groups. To limit the possibility that both types of products were exposed to dissimilar trends, we exclude discounters. With the same intent, we also remove all first-price products from the comparison group.[39] Indeed, first-price products are distant substitutes for national brands and then compete for different segments of the demand. More importantly, they are likely to be affected differently by costs shocks, because these products are characterized by higher pass-through than NB products.

Second, given the large number of product categories available in our database, we remove from the sample all product categories for which the national brand and private label trends do not satisfy the parallel trend assumption in the pre-LME period. To do so, we borrow from the insights of previous studies (see, *e.g.*, Weinberg and Hosken, 2013; Hosken *et al.*, 2018), and test for each product category whether national brands have a specific linear trend compared to private labels during the pre-LME period. We thus estimate the following OLS regression using pre-LME data only:

$$\ln\left(P_{ikt}\right) = \alpha Month_t + \beta Month_t \times T_{ik} + \delta T_{ik} + \mu_{ik} + \varepsilon_{ikt} \tag{10}$$

where P_{ikt} denotes the monthly average price of a chain-product pair ik at month t, $Month_t$ indicates the monthly period, T_{ik} is a dummy variable that takes the value one when the chain-product pair ik is a national brand, and μ_{ik} are a set of chain-product fixed effects.[40] We then remove all product categories for which the β coefficient is statistically different from zero at the 5% significance level. Note that imposing this selection criterion substantially reduces the sample size and the scope of categories covered.[41] We provide in Appendix B.3 the list of product categories that satisfy the parallel trend assumption.

[39]We include PL-D and PL-FP in the comparison group as a robustness check in Section 7.1.

[40]Note that throughout the empirical sections, "chain-products" stand for "products" in the theory section.

[41]This excludes from the analysis categories such as bottled water or potato chips, for instance.

We also discuss the underlying hypotheses of the selection procedure and perform several robustness tests using alternative selection procedures in the Online Appendix (Section C). Overall, we show that using alternative methods of selection – and varying the size and the composition of the final sample – does not alter our results.

Finally, focusing on the pre-LME period, we aggregate all product categories of our final sample and weight observations by their expenditure shares, and check that, overall, the prices of NB and PL products have evolved identically. As expected, there is no significant difference in the trends of the two types of brands (estimated coefficient: 0.0001 (0.0002)). This result thus supports the parallel trends assumption in the pre-LME period, a prerequisite to conduct a DiD analysis. Furthermore, we observe that the trends in prices of NB and PL products are not significantly different from zero in the pre-LME period.[42]

5.1.4. Summary Statistics on the Final Sample

Table 2 reports the composition of the final sample, split into affected and comparison groups. Despite our meticulous selection process, we keep a very large sample composed of 26,656 products aggregated into 76 product categories and 27 families of products. Overall, the monthly price data of the products retained in the final sample are calculated based on about 10 millions purchase observations. The most purchased product family is dairy products (24.19% of transactions), followed by non-alcoholic beverages (16.91%), sauces (7.51%), confectionery products (6.79%), and poultry products (5.82%). In terms of expenditures, the final sample represents more than 30 million euros of cumulative food expenditures over the period 2006-2010, which represents about 20% of total purchase expenditures on NB and PL products recorded in the KWP survey. Note that, by construction, the affected group contains only national brands, and the comparison group only private labels. As expected, we observe that affected products are more expensive than comparison products. The average monthly unit price of NB products is 10.53, while it is 8.22 euros for PL products. The allocation of expenditures between the affected and comparison groups is representative of the relative market shares of NB and PL in France during this time period. Indeed, national brands account for almost three quarters (72.93%) of the households' expenditures in our final sample, the rest being spent on private labels.

[42]We obtain a point estimate of the trend coefficient of 0.0001 for both national brands and private labels with *p-values* of 0.38 and 0.93, respectively.

Table 2: Summary Statistics for Affected and Comparison Groups

	Affected group	Comparison group	Total
Panel A: Product			
Number of products	17,744	8,912	26,656
Number of product categories	76	76	76
Number of product families	27	27	27
Average number of products per category	233.47	117.26	350.74
Number of chain stores	77	69	86
Panel B: Brand type			
Percentage of NB products	100	–	66.57
Percentage of PL products	–	100	33.43
Panel C: Price			
Mean of monthly average product price	10.53	8.22	10.02
S.D. of monthly average product price	23.60	11.72	21.49
Min. of monthly average product price	0.01	0.07	0.01
Max. of monthly average product price	3151.15	1878.23	3151.15
Panel D: Purchase transaction			
Number of purchase transactions	6,213,600	3,174,937	9,388,537
Total expenditures	23,101,467	8,583,887	31,685,354
% of KWP expenditures	19.41	20.22	19.62

Notes: The table reports summary statistics on the composition of the affected and comparison groups as well as for the final sample. Statistics are calculated over the pre- and post-LME periods. The last row (% of KWP expenditures) gives the share (in %) of total expenditures in the final sample relative to the total expenditures of the corresponding brand type in the KWP data.

5.2. Before and After Analysis

Before running our DiD analysis, we compare the evolution of (deflated) prices between the pre- and post-LME periods by running a simple time difference regression, to have a rough overview of the effect of the LME on prices. This time difference simply captures the specific price effect of the post-LME period, but not a hypothetical causal effect of the LME. The estimation results presented in column (1) of Table 3 are obtained by regressing the average monthly price of a chain-product pair against a dummy variable indicating the post-LME period and a set of chain-product fixed effects. We estimate that, after the LME, food retail prices have decreased on average by 1.3% relatively to the consumer price index.

In column (2), we interact the $PostLME$ variable with national brand and private label dummies. The point estimates indicate that the prices of national brands have decreased, on average, by 1.4% whereas the prices of private labels remained stable in comparison with the average prices observed in the pre-LME period. This suggests at first sight that the prices of private labels have not been affected by the law, a finding consistent with the theoretical predictions of scenario (i) in Table 1. In the next section, we test whether the difference in price changes between the two brand types are statistically significant.

Table 3: Price Changes around the LME

Dependent variable: (log) price (P_{ikt})		
Variable	(1)	(2)
PostLME	-0.0126	
	(0.0019)	
PostLME × PL		0.0024
		(0.0040)
PostLME × NB		-0.0136
		(0.0020)
Chain-product FE	Yes	Yes
R^2	0.989	0.989
Observations	1,919,906	1,919,906

Notes: The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. Standard errors, shown in parentheses, are clustered at the retail chain level.

5.3. The Price Effect of Authorizing Input Price Discrimination

We now estimate the direct price effect of the LME by comparing the mean change in prices that NB products have experienced between the pre- and post-LME periods, to the mean change in prices of PL products. In order to eliminate shocks that affect in a similar way both brand types in the pre- and post-LME periods, we run a DiD analysis by estimating the following OLS regression model:

$$\ln(P_{ikt}) = \beta T_{ik} \times PostLME_t + \delta T_{ik} + \gamma PostLME_t + \mu_{ik} + \varepsilon_{ikt} \tag{11}$$

where P_{ikt} denotes the monthly average price of a chain-product pair, $PostLME_t$ is a dummy variable equal to one for months following the introduction of the LME, and μ_{ik} is a set of chain-product fixed effects. The coefficient of interest β measures the average price effect of the LME and can be interpreted as the mean effect resulting from authorizing input price discrimination on the relative retail prices of national brands.

In column (1) of Table 4, we report the estimates of the baseline regression model, see equation (11), where observations are weighted by their expenditure shares in the pre-LME period. Standard errors are clustered to control for heteroskedasticity and serial correlation within chain. Compared to private labels, we observe that the price of NB products has significantly decreased by about 1.6%, on average, after the introduction of the LME. In order to control for unobserved shocks that could have differently affected the two groups of products, we augment our baseline specification by additional time-variant controls. In column (2), we first include chain-month fixed effects. For instance, if a retail chain selling more private labels than branded products decided to change its strategy by cutting down its prices, this would bias upward the estimated effect of the LME. Adding chain-month fixed effects, we observe that the point estimate does not significantly differ

Table 4: Authorizing Input Price Discrimination and Changes in Prices

Dependent variable	(log) price (P_{ikt})			(log) $\lvert\widehat{P}_{ik}^{post} - \widehat{P}_{ik}^{pre}\rvert$
		With monthly trend by		
	Baseline	Chain	Category	
	(1)	(2)	(3)	(4)
Treatment				0.3534
				(0.0357)
Treatment × PostLME	-0.0160	-0.0162	-0.0262	
	(0.0045)	(0.0047)	(0.0052)	
PostLME	0.0024			
	(0.0040)			
Chain-product FE	Yes	Yes	Yes	No
Chain-month FE	No	Yes	No	No
Category-month FE	No	No	Yes	No
Category FE	No	No	No	Yes
R^2	0.9886	0.9890	0.9893	0.4002
Observations	1,919,906	1,919,559	1,919,872	100,862

Notes: Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The point estimate of the *Treatment* variable is absorbed by the chain-product fixed effects in columns (1)-(3) and thus not available. Standard errors, shown in parentheses, are clustered at the chain level.

from the one obtained in column (1). In column (3), we instead add category-month fixed effects to control for demand and cost shocks at the category level. For instance, a shock on raw sugar prices would affect the sweetening products category that is over-represented in the national brand group, thereby introducing a potential bias. We obtain a point estimate that is almost twice that of column (1) which suggests that important shocks affecting some specific product categories biased our estimates. Controlling for these shocks, we find that authorizing input price discrimination has reduced by 2.6% the price of national brands as compared to private labels. This difference in the point estimates is not surprising since it is very likely that some product categories have experienced specific demand and/or cost shocks caused by contemporaneous events.[43] Henceforth, our preferred specification is the one adopted in column (3); it corresponds to the following regression:

$$\ln\left(P_{ikt}\right) = \beta T_{ik} \times PostLME_t + \delta T_{ik} + \gamma PostLME_t + \mu_{ik} + \eta_{ct} + \varepsilon_{ikt}, \qquad (12)$$

where η_{ct} represents a set of specific category-month fixed effects.

In all specifications, we highlight a negative coefficient, *i.e.* a relative price decrease of NB products. Finally, it is worth mentioning that this price effect is not sensitive to the date on which the LME effect is assumed to materialize (see the results of the robustness

[43] In Section 7.2, we explore the issue of confounding effects that may arise from two majors events of the late 2010s, *i.e.* the 2007-2008 food crisis and the Great Recession.

tests on the starting date of the LME in the Online Appendix A).

5.4. Discussion

We now discuss the interpretation of the β coefficient of equation (12) in light of our theoretical predictions reported in Table 1. The question is whether this coefficient can be interpreted as a measure of the causal effect of the LME on national brand prices. This is the case only if retailers have not changed the prices of private labels in reaction to the lift of the ban, a point that is not obvious as highlighted by our theoretical model.

Our before-and-after analysis highlights that PL products have not experienced any significant change in prices in the post-LME period compared to the pre-LME period (see column 2 in Table 3). This may indicate that PL products have not been impacted by the LME. Assuming that this is the case – which corresponds to scenario (i) presented in Table 1 – PL products constitute a valid counterfactual group. The β estimates of the DiD analysis reported in Table 4 can then be interpreted as the causal effect of the reform on the prices of NB products.

To have an idea of the magnitude of the estimated causal effect over all products in this scenario, we first do a back-of-the-envelop calculation, which indicates that the LME has caused a final price decrease of -1.89% on average.[44] We then relate the point estimate of β obtained in column (3) to the average monthly spending on national brands per household over the pre-LME period. Using the household panel and the entire set of transactions reported in the KWP, we calculate that households spend on average 175.46 € per month on national brands. Multiplying this value by the point estimate and assuming that households' shopping basket remained unchanged during the period, we calculate that authorizing input price discrimination has reduced by -4.60 € $(= -0.0262 \times 175.46)$ the average monthly price of the shopping basket of national brands in the pre-LME period compared to that of private labels, all else equal.[45]

Suppose instead that, as shown in the theoretical analysis, retailers have modified (up or down) the prices of PL products in response to a change in the input prices of NB products – see scenarios (ii)-(iv) in Table 1 – but that concomitant symmetric shocks may offset this price change in the before-and-after analysis. In that case, the β estimate of the DiD analysis reported in Table 4 can only be interpreted as the relative effect of the LME on national brand prices as compared to private label prices.

[44] To compute this back-of-the-envelope percentage, we used the mean of monthly average prices of national brands (10.53) and private labels (8.22) and the percentage of observations per brand type (67/33) available in Table 2.

[45] Taking into account the structure of expenses on NB and PL products in the pre-LME period, this amounts to a decrease of 1.90% in the average monthly expenditure by household.

We consider in turn scenarios (ii), (iii) and (iv) from Table 1. Scenario (iv) can be quickly dismissed as it would lead to a positive β. Indeed, we know from Lemma 1 that, in this case, the prices of PL products would increase less than the prices of NB products. In contrast, the two scenarios (ii) and (iii), which share the common prediction that the law causes a decrease in the price of NB products relatively to PL products, are consistent with the sign of the β estimate. However, in these scenarios the estimated coefficient β either underestimates or overestimates the causal effect of the LME on national brand prices. Indeed, in scenario (ii), all prices decrease, hence the β coefficient underestimates the overall effect of the ban on the prices of NB products. In contrast, in scenario (iii), as the prices of PL products increase, the β overestimates the effect of the ban on prices of NB products. Note however that in theory, when the input price variations are limited, the effect of the ban must be larger on NB than on PL products. This is confirmed empirically by the estimated coefficient displayed in column (4) of Table 4.[46] Moreover, the share of NB products in the households' expenditures and their absolute mean price value are larger than those of PL products. Therefore, we are confident that the estimated β gives the true *direction* of the overall effect of the ban on the price of the households' average shopping basket in the two scenarios.

Finally, the empirical analysis could be consistent with a last case (not predicted by theory), in which all prices would increase but prices of PL products would increase to a larger extent than that of NB products. Such a case would also lead to a negative β coefficient. Column (4) of Table 4 however dismisses this scenario.

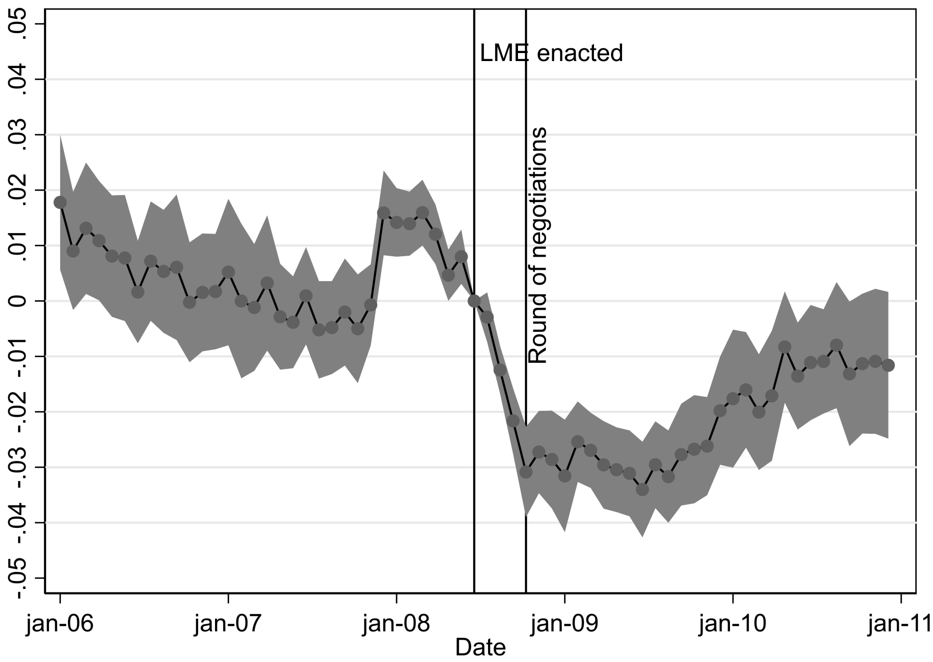
This discussion, prompted by the back-and-forth between theory and empirics, enables us to revisit the use of private labels or rivals' products as a control group in a quasi-experimental approach, which is rather usual in the retrospective merger literature (see, *e.g.*, Ashenfelter and Hosken, 2010; Ashenfelter *et al.*, 2013; Weinberg and Hosken, 2013; Miller and Weinberg, 2017; Björnerstedt and Verboven, 2015). These articles usually refer to the Deneckere and Davidson (1985) simplified framework of Bertrand competition, under which a merger between producers is likely to increase the prices of the products of the merging firms and, in reaction, but to a lower extent, the prices of their rivals' products (including that of private labels). In this framework, the obtained coefficient can therefore be considered as a lower bound of the causal effect of the merger. However, due to data availability issues, these articles use retail prices instead of input prices. Indirectly, they

[46] We first compute the average price of a given chain-product pair ik over the pre- and post-LME periods. Then, we determine the change in prices, in absolute value, between these two periods and we regress this (log) difference against the dummy variable T_{ik} as follows:

$$\ln\left[|\widehat{P}_{ik}^{post} - \widehat{P}_{ik}^{before}|\right] = \delta T_{ik} + \eta_c + \varepsilon_{ik}$$

where η_c corresponds to a set of product-category fixed-effects.

Figure 2: Event Study of the Effect of the LME on Log Prices



Notes: The figure plots the estimated specific deviation of the change in prices of national brands between month l and the reference month ($l = 0$) relative to the change in prices observed for private labels between these two months. The reference month corresponds to the month before the introduction of the LME (*i.e.*, July 2008). The shaded area represents the confidence interval of the point estimate at the 5% significance level.

assume that retailers do not distort the price effect, but this is far to be guaranteed with multi-product retailers and/or non-linear contracts. As in our model, the effect of a price increase of the merging firms' in-store product on the price of non merging firms' products (private labels or any rival products) is ambiguous, which challenges their use as a control group in an upstream merger analysis.

6. Extensions

This section further explores the heterogeneity of the price effect of the LME along several dimensions: over time, across products and retail groups.

6.1. The Price Effect over Time

The authorization to price discriminate came into force on August 5, 2008, for immediate application. Some retailers, whose contracts included a price review clause, thus seized the opportunity to renegotiate the term of sales of their 2008 contracts, without waiting for the opening of the next annual round of negotiations.

In order to explore how the price effect of the law unfolded over time, we adopt an "event study" approach by decomposing by month the change in prices of NB products between the pre- and post-LME period relative to that of PL products. To do so, we estimate the following OLS regression model:

$$ln\left(P_{ikt}\right) = \sum_{l=-32}^{32} \beta_l T_{ik} \times Month_t^l + \delta T_{ik} + \sum_{l=-32}^{32} \gamma_l Month_t^l + \mu_{ik} + \eta_{ct} + \varepsilon_{ikt}$$

where $Month_t^l$ is a dummy variable equal to one for the l-th month relative to July 2008, the month preceding the passing of the LME and which is chosen as reference. As before, we control for chain-product pair specific effects μ_{ik} and category-month specific effects η_{ct}. The coefficient β_l measures the average specific deviation of national brand prices relative to the price variation of private labels between the reference month ($l = 0$) and month l.

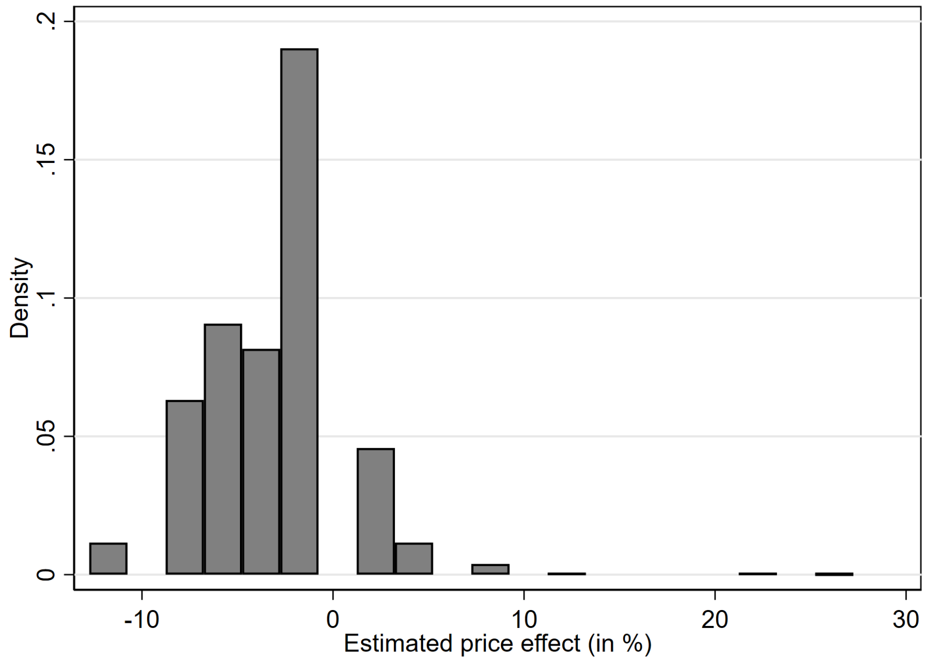
The estimated coefficients $\widehat{\beta}_l$ are plotted in Figure 2. A large number of the coefficients has a point estimate close to zero in the pre-LME period, which is consistent with the assumption of parallel price trends at the aggregate level when controlling for specific category price trends. We observe however a short-lived peak from December 2007 to March 2008, at the heart of the food crisis, where the prices of NB products increase relatively more than those of PL products (by about 1.4%). These price differences could inflate our measure of the causal effect in the DiD analysis. However, the event study takes July 2008, which is after the peak, as the reference month, and still exhibits a drop within an interval of (-3.40%, -0.8%) in the prices of NB relative to PL products in the months following the LME from November 2008 to December 2010. The price drop has reached its maximum in June 2009, but the change in prices remained stable – around -3% – from November 2008 until the end of 2009. This result suggests that a large part of the price decrease took place as soon as the law came into force. In the long term, however, we observe that the price effect is lessening, reaches a minimum of -0.8% in September 2010, and is on average about -1.29% over the year 2010.

6.2. Heterogeneous Price Effect Across Products

The intensity of producers' competition and the substitution between national brands and private labels vary across product categories. We thus expect that the effect of authorizing input price discrimination will be heterogeneous across product categories. We also analyze how this effect varies according to the initial price gap between national brands and private labels, a proxy for the product competition inside the category.

Effect by Product Category. We estimate equation (12) for each product category separately, replacing the category-month fixed effects by monthly dummies. Figure 3

Figure 3: Distribution of Estimated Price Effects by Product Category



Notes: The graph corresponds to the distribution of price effects estimated by product category. The product categories with a non-significant point estimate are withdrawn from the analysis. Observations are weighted by the expenditure shares of product categories, calculated at the national level during the pre-LME period.

presents graphically the distribution of the estimated price effects and shows that 72% of product categories have experienced a price decrease, which reinforces the robustness of previous findings. The estimated price effects range from -12.76% to +26.27%. The three largest decreases are observed for *other poultry w/o EAN* (-12.76%), *chocolate bars* (-11.02%), and *frozen snails* (-8.39%), whereas *whole turkey w/o EAN* (+26.27%), *frozen fruits, juice and puree* (+22.60%) and *freeze-dried vegetables* (+11.69%) have known significant price increases. All these product categories, except chocolate bars, are however marginal in terms of sales, and the bulk of products have experienced a fall in prices between 0 and 10% due to the passing of the LME.

Effects by Initial Price Gap between National Brands and Private Labels. We first compute, for each product category, the gap between the average price of NB and PL products in the pre-LME period. Although an imperfect proxy, this gap may indeed reflect the competitive pressure between NB and PL products within a category. We split the distribution of the price gaps in three classes (0 to 20%, 20% to 80%, and 80% to 100%). Table 5 then presents the effect of the LME for each of these categories. We highlight that the decrease in prices of NB relative to PL products is larger when the competitive pressure exerted in the product category is initially soft (for classes 20%-80% and 80%-100%).

Table 5: Price Gap between NB and PL products

Dependent variable: (log) price (P_{ijt})		
	(1)	(2)
Treatment × PostLME	-0.0262	
	(0.0052)	
Treatment × PostLME × Price Positioning 0-20		-0.0113
		(0.0064)
Treatment × PostLME × Price Positioning 20-80		-0.0261
		(0.0055)
Treatment × PostLME × Price Positioning 80-100		-0.0310
		(0.0045)
Chain-product FE	Yes	Yes
Category-month FE	Yes	Yes
R^2	0.989	0.989
Observations	1,919,872	1,919,872

Notes: Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The point estimate of the *Treatment* variable is absorbed by the chain-product fixed effects in columns (1)-(2) and thus not available. Standard errors, shown in parentheses, are clustered at the chain level.

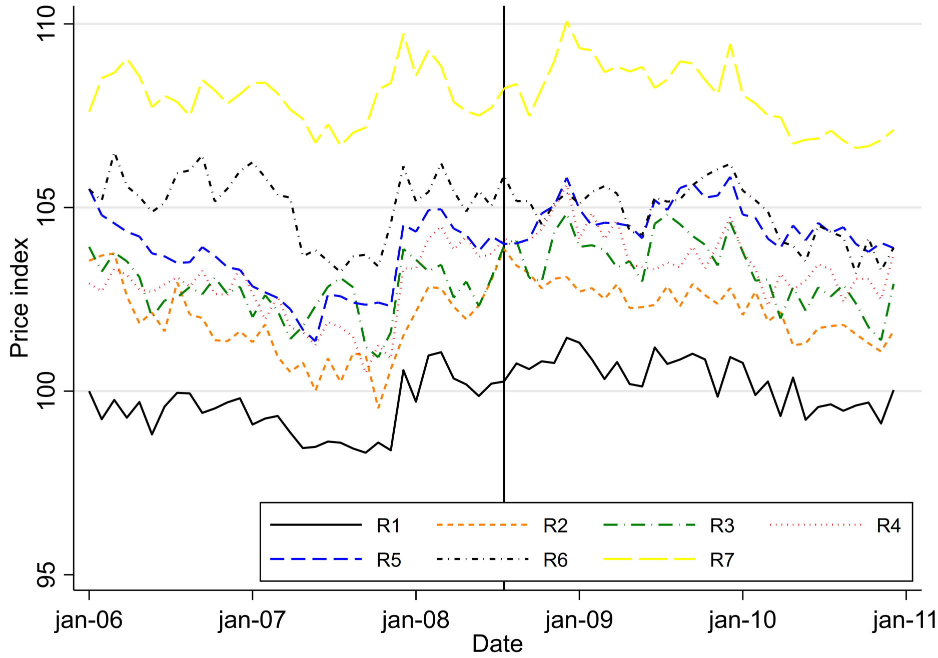
6.3. Heterogeneous Price Effects across Retail Groups

Although our model does not consider heterogeneous retailers for the sake of simplicity, it is likely that their price reactions differed according to their bargaining power, pricing strategies, or the competitive environment, for instance. In this section, we explore whether retailers reacted differently to the LME reform. To do so, we conduct the analysis at the retail group level, *i.e.* the level where negotiations with producers take place. We gather the chains owned by each of the seven major retail groups, denoted R_1 to R_7, and constitute an eighth group embedding all the remaining small chains.

We first construct a monthly national brand price index by retailer to study the price positioning of each retailer over time. To do so, we regress the log of prices of NB products included in the final sample conditional on product fixed-effects and retail group-month fixed effects. We then normalize retailer-month fixed effects by taking the first month of R_1 as base 100. We plot the monthly price indices in Figure 4. We observe that the price ranking of the groups is almost unchanged over the period of study. Interestingly, we note that the price dispersion seems to have decreased in the post-LME period. To verify this pattern, we compute for each month the standard deviation of the price indices, and compare the average standard deviation in the pre-LME period to that in the post-LME period. The T-test confirms a significant decrease in the price dispersion in the post-LME period (T-test: p<0.0001).

In Table 6, we report the estimation results of equation 12 when the average treatment effect is split by retailer. We rank retailers from R_1 to R_8 according to the level of their

Figure 4: Price Index of National Brands by Retail Group



Notes: The graph plots the monthly price index of NB products by retail group over the period of study. The first month of the cheapest retail group, R_1, is taken as base 100. We do not plot R_8's price index as it gathers heterogeneous retail groups.

mean prices in the pre-LME period (R_1 being the less expensive one). In column (1), we test whether the less and most expensive retailers (among the major retailers) experience different price changes compared to other retailers. We find that only the most expensive retailer, R_7, has significantly decreased its prices with respect to other retailers. Column (2) decomposes the effect by retail group. While all retailers have dropped the prices of their NB products by at least 2%, we find that the most expensive retailers in the pre-LME period, *i.e.* R_6 to R_8, have experienced a larger decrease, by slightly more than 3%, whereas the least expensive ones have experienced a decrease by less than 2.5%. This result suggests that the lift of the ban has forced high-price retailers to cut down drastically the prices of NB products in order to maintain their market share. This differentiated price effect has then contributed to the price convergence of NB products after the implementation of the law.

7. Robustness

We test the robustness of our main result to alternative definitions of the comparison group (Section 7.1) and to potential confounders, specifically the 2007-2008 global food crisis and the Great Recession (Section 7.2).

Table 6: Estimated Price Effect by Retailer

Dependent variable: (log) price (P_{kit})		
	(1)	(2)
Treatment × PostLME	-0.0250	
	(0.0055)	
Treatment × PostLME × R1	0.0021	-0.0228
	(0.0017)	(0.0053)
Treatment × PostLME × R2		-0.0250
		(0.0053)
Treatment × PostLME × R3		-0.0210
		(0.0054)
Treatment × PostLME × R4		-0.0187
		(0.0053)
Treatment × PostLME × R5		-0.0256
		(0.0054)
Treatment × PostLME × R6		-0.0305
		(0.0053)
Treatment × PostLME × R7	-0.0076	-0.0327
	(0.0018)	(0.0054)
Treatment × PostLME × R8		-0.0395
		(0.0065)
Chain-product FE	Yes	Yes
Category-month FE	Yes	Yes
R^2	0.989	0.989
Observations	1,919,872	1,919,872

Notes: Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. Standard errors, shown in parentheses, are clustered at the retail group level.

7.1. Alternative Definition of the Comparison Group

We construct four alternative definitions of the comparison group by considering all the types of private labels: private labels sold by traditional retailers (PL) and by discounters (PL-D), as well as first-price private labels (PL-FP). For each comparison group, the product sample is redefined to guarantee the parallel trend assumption. This implies that the number of observations, products and product categories vary with the definition of the comparison groups.

The first row in Table 7 reports our baseline results with an average effect of 2.62%. The variant of the comparison group considered are less likely to be exposed to the *Edgeworth-Salinger effect*. Indeed, first price PL products (PL-FP) are often sold at marginal cost, and private labels sold at discounters (PL-D) are only indirectly exposed to the *Edgeworth-Salinger effect* because discounters offer few NB products.[47] Therefore, in the four considered alternative definitions, the effect of the law is less likely to be

[47] This is the case, in particular, when including the PL products of discounters that barely coexist with NB products in retailers' shelves. However, the prices of PL-D may still react to the potential price change of PL products sold by traditional retailers.

Table 7: Alternative Definitions of the Comparison Group

	$\widehat{\beta}$				# of	% of KWP
Comparison group	Coef.	S. E.	Obs.	R^2	Cat.	expend.
Baseline (PL)	-0.0262	0.0052	1,919,872	0.9893	76	19.62
PL & PL-FP	-0.0182	0.0039	2,143,344	0.9836	77	21.48
PL & PL-D	-0.0126	0.0041	1,796,503	0.9893	69	17.85
PL, PL-FP & PL-D	-0.0150	0.0035	2,140,553	0.9834	75	20.74
PL-D	-0.0171	0.0077	1,679,180	0.9816	60	20.46

Notes: The table gives the point estimate of the β coefficient using various comparison groups. Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. Standard errors (denoted S. E.) are clustered at the chain level. The last two columns report the number of categories retained in the final sample and the share (in %) of total expenditures relative to the total expenditures of the corresponding brand type in the KWP data, respectively.

overestimated. Interestingly, in all the four cases we find a smaller effect that ranges from -1.26% to -1.82%. Note however that, as explained in Section 5.1.3, these control groups are more likely to be exposed to different demand and supply shocks, which make them less relevant than our baseline comparison group.

7.2. Potential Confounders

The years around the passing of the LME have witnessed two events that have significantly impacted the supply and demand of food products worldwide: the “2007-2008 global food crisis”, and the 2008 Great Recession. In this section, we conduct robustness tests to assess the sensitivity of our main result to these potential confounders.

First, the end of year 2007 was marked by a sudden surge in prices, reaching a spike in early 2008 (February-May), that contrasts with the deflationary trend observed in the previous months. This abrupt change in food prices resulted from the “2007-08 global food crisis”, which significantly impacted the prices of major agricultural commodities (maize, rice, soybeans, wheat).[48] Producers and retailers were constrained to pass part of this increase through to consumers. This shock may have affected differently product categories with heterogeneous relative contribution to the affected and comparison groups, thus introducing a potential bias in our estimate. To limit this concern, we have eliminated first-price products from our sample because these products usually exhibit a higher pass-through than products sold with a higher margin in the same category. Moreover, by selecting a sample of products that satisfy the parallel trend assumption, we also made sure to keep products for which national brands and private labels prices had evolved

[48] The food crisis was due to the conjunction of several factors. On the one hand, demand for agricultural commodities surged due to the growing development of biofuels and the increasing demand of protein diets in some parts of the world. On the other hand, in 2007 supply was impacted by transitory phenomena (such as climate events and a fall in world-food stockpiles, among others). Overall, the increased tension between supply and demand led to a historic rise in prices.

similarly during the food crisis. This careful selection process has limited the impact of the food crisis on our results. However, to test further the robustness of our results, we run an alternative estimation removing the whole period of the food crisis (*i.e.*, from September 2007 to September 2008). We obtain a point estimate of -0.0253 that is still statistically significant at the 1% level and close to our baseline coefficient (-0.0262). This tends to demonstrate that our evaluation of the LME effect is not sensitive to this particular event.

Second, our period of study is also concomitant with the Great Recession. According to the NBER, this global crisis started in the US in December 2007, and the bulk of its effect lasted until June 2009. In many countries, the Great Recession caused a negative income shock for households. Several studies, such as Griffith *et al.* (2016) for the UK and Nevo and Wong (2019) for the US, have documented that households have significantly decreased their (real) food expenditures after the crisis by modifying their shopping habits.[49] Such changes in consumers' demand could lead retailers to increase the relative price of private labels. This would challenge our identification assumption that the prices of NB and PL products would not have been affected but for the LME. In what follows, we therefore investigate whether the Great Recession has affected the relative market share of private labels.[50]

In Table 8, we present the OLS estimation results of the relative market share of private labels on the time span of the Great Recession. The estimation is conducted at the product category level. All regressions control for product-category fixed effects and seasonal fixed-effects. In column (1), we consider that the effect of the Great Recession covers a wide period starting in December 2007 to December 2010. The result shows no significant effect. In columns (2) and (3), we follow the approach of Dubé *et al.* (2018) and decompose the dummy variable into two sub-periods, the first one corresponding to the Great Recession and the following corresponding to a post-recession period. In column (2), in which the Great Recession period spans from December 2007 to June 2009 as reported by the NBER, we find no significant effect during or after the recession. As there is evidence that the effects of the Great Recession have been delayed in France as compared

[49] For instance, Griffith *et al.* (2016) highlight that (real) food expenditures fell by 6% between the pre-recession (2005-07) and the post-recession (2010-12) years in the UK. They document that UK households reacted to budget cuts by reducing the price of their shopping basket while keeping their calorie intake and the nutritional quality of the food unchanged. To do so, they visited more stores, searched for better deals, switched for cheaper retailers and brand types (*e.g.* private labels), and substituted away from expensive nutrients. Nevo and Wong (2019) show further that US households reacted to the drop in the food budget by allocating more time to home production at the expense of shopping.

[50] Unlike in other developed countries (see for instance Griffith *et al.* (2018) for the UK), the market share of private labels has reached a steady state in France at the end of the 2000s. This is confirmed by running a series of OLS regressions with various sets of fixed-effects and different levels of clustering. All estimation results reject the existence of a (linear) trend in the pre-LME period, that is, between January 2006 and August 2008.

Table 8: Changes in relative Market Shares of PL vs. NB during the Great Recession

Dependent variable: relative market share of private label by product category			
	Recession (Post Dec 07) (1)	Recession (Dec 07 - Jun 09) (2)	Recession (Apr 08 - Jun 09) (3)
Post-December 07	0.0427		
	(0.0387)		
Recession		0.0043	0.0141
		(0.0311)	(0.0276)
Post-recession		0.0773	0.0801*
		(0.0479)	(0.0450)
Category FE	Yes	Yes	Yes
Seasonal FE	Yes	Yes	Yes
R^2	0.7817	0.7828	0.7828
Observations	4261	4261	4261

Notes: The dummy variable Great Recession takes value one from December 2007 onward. Observations are weighted by the expenditure shares of product categories, calculated at the national level during the pre-LME period. Standard errors are clustered at the product category level.

to the US, in column (3) we use the period retained by the French Economic Association for the Great Recession in France (April 2008 to June 2009).[51] Column (3) reports only a barely significant effect of the post-recession period on the relative market shares of private labels. As a result, we are confident that the relative prices of NB products was not affected by potential changes in consumers' demand driven by the Great Recession.[52]

8. Conclusion

This paper investigates, both theoretically and empirically, how input price discrimination affects final prices. We first contribute to the theoretical literature by analyzing input price discrimination in a setup featuring upstream competition between the producers of national brands and of private labels, and downstream competition between multi-product retailers. In a secret contracting environment, we highlight that, in addition to the *competition effect* usually pointed out in the literature, two new economic forces arise due to the multi-product activity of retailers, namely, the *bargaining leverage effect*, and the *Edgeworth-Salinger effect*. The interactions of these three effects result in scenarios which all predict a differentiated impact of input price discrimination on national brands and private labels prices. A reform that lifted the ban on input price discrimination in

[51] The French Economic Association (AFSE) uses the same dating procedure as the NBER to determine the business cycles in France (See https://www.afse.fr/fr/cycles-eco/dates-500216).

[52] These results are in line with those of Dubé *et al.* (2018) who have shown that the negative effect of the Great Recession on income and wealth only had a small impact on the market share of private labels in the US. Moreover, there is evidence that France was among the OECD countries that were least affected (see, for instance André *et al.* (2015)).

France in 2008 provides a quasi-natural experiment that allows us to determine which of these scenarios are the most plausible. Leveraging on our theoretical predictions, we select private labels as the comparison group and conduct a DiD analysis of the effect of the LME reform. Our results show that the lift of the ban on input price discrimination caused a significant decrease in the relative prices of national brands by 2.62%. We explore the heterogeneity of the effect of this law along several dimensions (by product category, by retailer, over time, ...).

Our article has important policy implications as it provides, to our knowledge, the first *ex-post* evaluation of the effect of input price discrimination on retail prices on thousands of food products that cover a large share of food expenditures. In particular, we highlight that, in the context of our study, the *competition effect* dominates other potential effects, such as the *bargaining leverage effect*, or the *exclusion effect*, which is often put forward by policy makers to advocate a ban on input price discrimination. In terms of policy evaluation, our article thus supports the reform at stake in our data, and shows that it reached its objective of causing a decrease in food prices. Another key contribution is to show that as the *Edgeworth-Salinger effect* holds in a retail competition framework, the choice of private labels as a comparison group is not neutral. Even if their input price is not directly affected by the lift of the ban, their coexistence with national brands on retailers' shelves is sufficient to make their retail price also react to the lift – in any direction. More broadly, this result also applies to *ex-post* evaluations of upstream mergers, in which rivals' retail prices are often used as a control group thus potentially leading to an upper bound estimate of the merger effect instead of a lower bound, as often argued in the literature.

Finally, providing a welfare analysis is beyond the scope of this article. To do so, a possible approach would be to focus on a handful of product categories and estimate a structural model of demand and supply. We leave this for further research.

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A. The Model

A.1. Well-Behaved Reaction Functions

Assumption 2 *(i)* ensures that the retailers' profit functions are concave in prices. The first order conditions given by equations (2) thus characterize the best response final prices $(p^r_{Ai}(p_{Aj}, p_{Bj}), p^r_{Bi}(p_{Aj}, p_{Bj}))$ of retailer $i \in \{1, 2\}$ who sells goods $k \in \{A, B\}$ as a (differentiable) function of the final prices of the rival retailer and of the input prices w_{Ai} and w_{Bi}.

Totally differentiating the FOC wrt. p_{Aj} and simplifying yields:

$$\begin{aligned}
\pi^i_{11}(.)\frac{dp^r_{Ai}}{dp_{Aj}} + \pi^i_{12}(.)\frac{dp^r_{Bi}}{dp_{Aj}} + \pi^i_{13}(.) &= 0,\\
\pi^i_{21}(.)\frac{dp^r_{Ai}}{dp_{Aj}} + \pi^i_{22}(.)\frac{dp^r_{Bi}}{dp_{Aj}} + \pi^i_{23}(.) &= 0.
\end{aligned}$$

Hence we obtain: $\frac{dp^r_{Ai}}{dp_{Aj}} = \frac{\pi^i_{12}(.)\pi^i_{23}(.) - \pi^i_{22}(.)\pi^i_{13}(.)}{\Delta(.)}$.

Similarly, totally differentiating the FOC wrt. p_{Bj} yields

$$\frac{dp^r_{Bi}}{dp_{Bj}} = \frac{\pi^i_{21}(.)\pi^i_{14}(.) - \pi^i_{11}(.)\pi^i_{24}(.)}{\Delta(.)},$$

where $\Delta(.) \equiv \pi^i_{11}(.)\pi^i_{22}(.) - \pi^i_{12}(.)\pi^i_{21}(.)$ is the determinant of retailer i's profit function, which is positive under Assumption 2 (i). Furthermore, numerators are positive under Assumption 2 (i) and (ii). Hence under Assumption 2 *(i)* and *(ii)*, $\frac{dp^r_{Ai}}{dp_{Aj}} > 0$ and $\frac{dp^r_{Bi}}{dp_{Bj}} > 0$: the prices of good $k \in \{A, B\}$ at the two retailers are strategic complements.

Finally, we consider the variation of the best response prices with respect to retailer i's unit costs w_{Ai} and w_{Bi}. Totally differentiating the two first order conditions with respect to w_{Ai} and to w_{Bi} for a given (p_{Aj}, p_{Bj}) yields the best-response pass-through:

$$\frac{dp^r_{Ai}(p_{Aj},p_{Bj})}{dw_{Ai}} = \frac{\pi^i_{22}D_1 - \pi^i_{12}D_2}{\Delta} > 0, \quad \frac{dp^r_{Ai}(p_{Aj},p_{Bj})}{dw_{Bi}} = \frac{\pi^i_{22}D_2 - \pi^i_{12}D_1}{\Delta}, \tag{A-1}$$

$$\frac{dp^r_{Bi}(p_{Aj},p_{Bj})}{dw_{Bi}} = \frac{\pi^i_{11}D_1 - \pi^i_{21}D_2}{\Delta} > 0, \quad \frac{dp^r_{Bi}(p_{Aj},p_{Bj})}{dw_{Ai}} = \frac{\pi^i_{11}D_2 - \pi^i_{21}D_1}{\Delta}. \tag{A-2}$$

Note that Assumptions 1 and 2 *(i)* are sufficient to ensure that the best response price for each good increases with its own unit cost.

A.2. Properties of the Linear Demand Specification

We consider the following demand function, that derives from Singh and Vives (1984) [53]

$$q_{ki} = \frac{1-p_{ki}-b(1-p_{kj})-a(1-p_{li}-b+bp_{lj})}{(1-a^2)(1-b^2)},$$

where $a \in [0,1]$ and $b \in [0,1]$. Note that a and b represent the degree of inter- and intra-brand competition, while the substitutability between products ki and lj is a combination between intra- and inter-brand substitution. The underlying assumption is that a representative consumer has a quadratic utility function and a budget of 1: $U(q) = \sum_{k,i} q_{ki} - \frac{1}{2}\sum_{k,i} q_{ki}^2 - a\sum_i q_{Ai}q_{Bi} - b\sum_k q_{k1}q_{k2} - a.b\sum_k q_{k1}q_{l2}$.

With this specification, it is straightforward that Assumptions 1 to 4 are satisfied. We derive the equilibrium final price

$$p^*_{ki}(w_{ki}, w_{kj}, w_{li}, w_{lj}) = \frac{2-b(1+b)+2w_{ki}+bw_{kj}}{4-b^2}.$$

A.3. Proof of Proposition 1

Assume a symmetric distribution of input prices across retailers: $w_{A1} = w_{A2} = w_A$ and $w_{B1} = w_{B2} = w_B$. By totally differentiating the first order condition (2), for $i = 1$ say, with respect to w_A, we obtain:

$$\begin{aligned}
\frac{dp^*_{A1}}{dw_A} &= -\frac{-\frac{\partial D_{A1}}{\partial p_{A1}} + \pi^i_{12}\frac{dp_{B1}}{dw_A} + \pi^i_{13}\frac{dp_{A2}}{dw_A} + \pi^i_{14}\frac{dp_{B2}}{dw_A}}{\pi^i_{11}}, \\
\frac{dp^*_{B1}}{dw_A} &= -\frac{-\frac{\partial D_{A1}}{\partial p_{B1}} + \pi^i_{21}\frac{dp_{A1}}{dw_A} + \pi^i_{23}\frac{dp_{A2}}{dw_A} + \pi^i_{24}\frac{dp_{B2}}{dw_A}}{\pi^i_{22}}.
\end{aligned}$$

Symmetry across retailers implies that $\frac{dp^*_{A1}}{dw_A} = \frac{dp^*_{A2}}{dw_A}$ and $\frac{dp^*_{B1}}{dw_A} = \frac{dp^*_{B2}}{dw_A}$. Solving this system of equations yields:

$$\frac{dp^*_{Ai}}{dw_A} = \frac{X-W}{E-D}, \qquad \frac{dp^*_{Bi}}{dw_A} = \frac{Z-Y}{E-D},$$

where

$$\begin{aligned}
W &= D_2\left[\pi^i_{12} + \pi^i_{14}\right] > 0, & Z &= D_2\left[\pi^i_{13} + \pi^i_{11}\right] < 0, \\
X &= D_1\left[\pi^i_{24} + \pi^i_{22}\right] > 0, & D &= (\pi^i_{12} + \pi^i_{14})(\pi^i_{21} + \pi^i_{23}), \\
Y &= D_1\left[\pi^i_{21} + \pi^i_{23}\right] < 0, & E &= (\pi^i_{13} + \pi^i_{11})(\pi^i_{24} + \pi^i_{22}).
\end{aligned}$$

The signs of W, X, Y and Z derive from Assumptions 1 to 3.

[53] Choné and Linnemer (2020) argue that this demand system can be traced back to the works of Levitan and Shubik (1971).

Under Assumption 3, we have $E - D > 0$. Furthermore $X - W > 0$ under Assumptions 1 to 3, hence $\frac{dp^*_{Ai}}{dw_A} > 0$. In contrast, the sign of $Z - Y$, hence of $\frac{dp^*_{Bi}}{dw_A}$, is ambiguous since $|\pi^i_{13} + \pi^i_{11}| > \pi^i_{23} + \pi^i_{21} > 0$ under Assumption 3 whereas $|D_1| > D_2 > 0$ under Assumption 1.

Linear specification. Deriving the equilibrium prices obtained in Appendix A.2 with the linear specification, we have $\frac{dp^*_{Ai}}{dw_A} = \frac{1}{2-b} > 0$ and $\frac{dp^*_{Bi}}{dw_A} = 0$.

A.4. Proof of Lemma 1

Consider now the comparison between $\frac{dp^*_{Ai}}{dw_A}$ and $|\frac{dp^*_{Bi}}{dw_A}|$.

- Assume first that $\frac{dp^*_{Bi}}{dw_A} > 0$, *e.g.* $Z - Y > 0$. Under Assumptions 2 and 3, we have:

$$-(\pi^i_{24} + \pi^i_{22}) \geq \pi^i_{12} + \pi^i_{14} > 0, \text{ and}$$
$$0 > -(\pi^i_{21} + \pi^i_{23}) \geq \pi^i_{13} + \pi^i_{11}.$$

  Hence $-(\pi^i_{24}+\pi^i_{22}+\pi^i_{23}+\pi^i_{21}) \geq \pi^i_{14}+\pi^i_{12}+\pi^i_{13}+\pi^i_{11}$. Furthermore under Assumptions 2 and 3, we know that $-(\pi^i_{24} + \pi^i_{22} + \pi^i_{23} + \pi^i_{21}) > 0$. Using $-D_1 > D_2 > 0$, we thus obtain

$$D_1(\pi^i_{24} + \pi^i_{22} + \pi^i_{23} + \pi^i_{21}) \quad \geq \quad D_2(\pi^i_{14} + \pi^i_{12} + \pi^i_{13} + \pi^i_{11}).$$

  We have shown that $E - D > 0$, hence $\frac{dp^*_{Ai}}{dw_A} \geq |\frac{dp^*_{Bi}}{dw_A}|$.

- Assume now that $\frac{dp^*_{Bi}}{dw_A} < 0$, *e.g.* $Z - Y < 0$. Then we show that, at any equilibrium in which $w_A = w_B$, we have $\frac{dp^*_{Ai}}{dw_A} \geq |\frac{dp^*_{Bi}}{dw_A}|$.

  Under Assumption 1, we have $|D_1| > D_2$. Furthermore, $\pi^i_{12} = \pi^i_{21}$. At a symmetric price equilibrium in which $w_A = w_B$, by symmetry across products, we also have $\pi^i_{23} = \pi^i_{14}$, $\pi^i_{13} = \pi^i_{24}$ and $\pi^i_{11} = \pi^i_{22}$, hence $(|\pi^i_{24} + \pi^i_{22}| + \pi^i_{23} + \pi^i_{21}) = (|\pi^i_{13} + \pi^i_{11}| + \pi^i_{12} + \pi^i_{14})$. Therefore at a symmetric price equilibrium, under Assumption 2 we obtain

$$|D_1|(|\pi^i_{24} + \pi^i_{22}| + \pi^i_{23} + \pi^i_{21}) > D_2(|\pi^i_{13} + \pi^i_{11}| + \pi^i_{12} + \pi^i_{14})$$

  which always holds under Assumption 1. The same still holds at the neighborhood of a symmetric price equilibrium.

A.5. Proof of Proposition 2

Candidate equilibrium in the discrimination case.

- Consider first the offer made by supplier U_{Bi} to retailer R_i. The constraint in supplier U_{Bi}'s program (3) is binding. The FOC with respect to w_{Bi} is then:

$$\begin{aligned} 0 &= \tfrac{dp^r_{Bi}}{dw_{Bi}} D_{Bi}(.) + (p^r_{Bi} - c)\left(\tfrac{\partial D_{Bi}}{\partial p_{Ai}} \tfrac{dp^r_{Ai}}{dw_{Bi}} + \tfrac{\partial D_{Bi}}{\partial p_{Bi}} \tfrac{dp^r_{Bi}}{dw_{Bi}}\right) \\ &+ \tfrac{dp^r_{Ai}}{dw_{Bi}} D_{Ai}(.) + (p^r_{Ai} - w^d_{Ai})\left(\tfrac{\partial D_{Ai}}{\partial p_{Ai}} \tfrac{dp^r_{Ai}}{dw_{Bi}} + \tfrac{\partial D_{Ai}}{\partial p_{Bi}} \tfrac{dp^r_{Bi}}{dw_{Bi}}\right). \end{aligned}$$

Reintegrating the retailers' FOCs (2) and simplifying yields:

$$0 = (w_{Bi} - c)\left[\tfrac{\partial D_{Bi}}{\partial p_{Bi}} \tfrac{dp^r_{Bi}}{dw_{Bi}} + \tfrac{\partial D_{Bi}}{\partial p_{Ai}} \tfrac{dp^r_{Ai}}{dw_{Bi}}\right]. \tag{A-3}$$

This holds for both suppliers B_i, hence there exists a symmetric equilibrium such that input prices are $w^d_{Bi} = w^d_{Bj} = c$.

- Consider now the offer made by supplier U_A to retailer R_i. The FOC associated to program (5) is as follows:

$$\begin{aligned} 0 = \tfrac{dp^r_{Ai}}{dw_{Ai}}&[(p^r_{Ai} - c)\tfrac{\partial D_{Ai}}{\partial p_{Ai}} + (p^r_{Bi} - w^d_{Bi})\tfrac{\partial D_{Bi}}{\partial p_{Ai}} + (w^d_{Aj} - c)\tfrac{\partial D_{Aj}}{\partial p_{Ai}} + D_{Ai}] + \\ \tfrac{dp^r_{Bi}}{dw_{Ai}}&[(p^r_{Ai} - c)\tfrac{\partial D_{Ai}}{\partial p_{Bi}} + (p^r_{Bi} - w^d_{Bi})\tfrac{\partial D_{Bi}}{\partial p_{Bi}} + (w^d_{Aj} - c)\tfrac{\partial D_{Aj}}{\partial p_{Bi}} + D_{Bi}]. \end{aligned}$$

Reintegrating the retailer's FOCs (2) and simplifying yields:

$$\begin{aligned} 0 = \tfrac{dp^r_{Ai}}{dw_{Ai}}&[(w_{Ai} - c)\tfrac{\partial D_{Ai}}{\partial p_{Ai}} + (w^d_{Aj} - c)\tfrac{\partial D_{Aj}}{\partial p_{Ai}}] + \\ \tfrac{dp^r_{Bi}}{dw_{Ai}}&[(w_{Ai} - c)\tfrac{\partial D_{Ai}}{\partial p_{Bi}} + (w^d_{Aj} - c)\tfrac{\partial D_{Aj}}{\partial p_{Bi}}]. \end{aligned}$$

There exists a symmetric equilibrium such that $w^d_{A1} = w^d_{A2} = c$.

- We have shown that $w^d_{A1} = w^d_{A2} = w^d_{B1} = w^d_{B2} = c$ sustains an equilibrium.

Uniqueness. Focusing on equilibria that are symmetric across the retailers and across products (*i.e.*, with $w_{Ai} = w_{Bi} = w$ for $i = 1, 2$), equation (A-3) implies that an equilibrium with $w \neq c$ would exist only if, in the continuation equilibrium,

$$\left[\tfrac{\partial D_{Bi}}{\partial p_{Bi}} \tfrac{dp^r_{Bi}}{dw_{Bi}} + \tfrac{\partial D_{Bi}}{\partial p_{Ai}} \tfrac{dp^r_{Ai}}{dw_{Bi}}\right] = 0. \tag{A-4}$$

Yet we have shown above in Appendix A.1 that at a symmetric equilibrium $\frac{dp^r_{Bi}}{dw_{Bi}} > |\frac{dp^r_{Ai}}{dw_{Bi}}|$. Furthermore, Assumption 1 ensures that $|\frac{\partial D_{Bi}}{\partial p_{Bi}}| \geq \frac{\partial D_{Bi}}{\partial p_{Ai}} \geq 0$, hence there is no vector of input prices that satisfies A-4.

Linear specification. With the linear specification defined in Appendix A.2, the equilibrium input prices are $w^d_{Ai} = w^d_{Bi} = c$, and the equilibrium final prices are $p^*(c, c, c, c) = \frac{1-b+c}{2-b}$.

A.6. Proof of Proposition 4

Candidate equilibrium in the non discrimination case. From proposition 3, we have a unique equilibrium $w^{nd}_{B1} = w^{nd}_{B2} = c$ and this is known by all - in what follows, we will thus omit the arguments w_{B1} and w_{B2} when obvious. Consider now U_A's negotiation program with R_1 given by equation (7). Now, *status quo* profits $\overline{\pi}^1_B(w_A)$ depend on w_A. Indeed, when R_1 does not distribute good A, R_2 still sells product A and buy it at the same price w_A, and therefore, the demand for product B at retailer R_1 may still depend on w_A. The negotiation program with R_2 is symmetric and the participation constraints of the retailers are binding, therefore U_A's program boils down to maximizing:

$$\underset{w_A}{Max} \sum_{i=1,2} (p^*_{Ai}(w_A) - c)D_{Ai}(.) + \sum_{i=1,2} (p^*_{Bi}(w_A) - c)D_{Bi}(.) - \sum_{i=1,2} \overline{\pi}^i_B(w_A). \tag{A-5}$$

We compute below the value of the FOC in $w_A = w_B = c$. In that case, the downstream equilibrium price is symmetric across retailers and products and denoted by p^*. After reintegrating the downstream firms' FOCS given by (2), the first order condition of program (A-5) can be simplified as follows:

$$\begin{aligned} X &= \frac{dp^*_{Aj}}{dw_A} \underbrace{[(p^* - c)(\frac{\partial D^*_{Ai}}{\partial p_{Aj}} + \frac{\partial D^*_{Bi}}{\partial p_{Aj}}) - (p^3_{Bi} - c)\frac{\partial D^3_{Bi}}{\partial p_{Aj}}]}_{\equiv \phi} \\ &+ \frac{dp^*_{Bj}}{dw_A} \underbrace{[(p^* - c)(\frac{\partial D^*_{Ai}}{\partial p_{Bj}} + \frac{\partial D^*_{Bi}}{\partial p_{Bj}}) - (p^3_{Bi} - c)\frac{\partial D^3_{Bi}}{\partial p_{Bj}}]}_{\equiv \psi}. \end{aligned}$$

Under Assumption 4 we have $\phi > 0$. Note that, under Assumption 1, we also have $\phi > \psi$. We also use lemma 1 which highlights that at a symmetric equilibrium $|\frac{dp^*_{Bi}}{dw_A}| < |\frac{dp^*_{Ai}}{dw_A}|$.

- Assume first that $\frac{dp^*_{Bi}}{dw_A} \geq 0$.
  - If $\psi \geq 0$, we have $X > \frac{dp^*_{Bi}}{dw_A}(\phi + \psi) \geq 0$, hence the supplier's FOC is positive in $w_A = c$: the equilibrium is thus such that $w^{nd}_A > c$.
  - If $\psi < 0$, in equilibrium it is possible that $w^{nd}_A < c$, especially if $|\psi| \geq |\phi|$
- Assume now that $\frac{dp^*_{Bi}}{dw_A} < 0$.
  - Assume that $\psi \geq 0$, at $w_A = w_B = c$, we have $X > \frac{dp^*_{Ai}}{dw_A}(\phi - \psi) \geq 0$ and $w^{nd}_A > c$.
  - Assume that $\psi < 0$, then $X > \phi\frac{dp^*_{Ai}}{dw_A} > c$.

The sign of X characterizes the input price for product A: if $X > 0$, in equilibrium, $w_A^{nd} > c$ while if $X < 0$, in equilibrium $w_A^{nd} < c$. QED.

Linear demand specification. With the linear specification defined in Appendix A.2, we have $\phi = (p^* - c)\frac{b}{(1+a)(1-b^2)} \geq 0$ and $\psi = (p^* - c)\frac{-ab}{(1+a)(1-b^2)} \leq 0$. The equilibrium input price is $w_A^{nd} = c + \frac{(1-a)b(1-c)}{2} > c$ whereas $w_{Bi}^{nd} = c$. Finally the final prices are $p_{Ai}^{nd} = p^d + \frac{(1-a)b(1-c)}{2(2-b)} > p^d$ and $p_{Bi}^{nd} = p^d$.

B. Data

B.1. The KWP Data

We use household panel data, collected by KWP, to analyze the evolution of food prices in France over the 2006-2010 period (see Kantar Worldpanel, 2010). The data contains information on the daily food purchases made for home consumption, including online purchases, by the households enrolled in the panel. After each shopping trip, a household registers the EAN, a 13-digit barcode – or a description of the purchase if a EAN is not available–, the quantity and the expenditure of each purchase by means of a handheld scanner. It also provides information on the location of purchase, *i.e.* the store type (*e.g.*, supermarket, specialized store) and the retail chain name. KWP does not provide the EANs but gives instead its own identifier. Each product is also described by a large set of attributes, the brand name and the name of the manufacturer. Products are clustered into more than 349 categories of food products, which are in turn aggregated into 61 families of products.

The composition of the panel varies over time but it is representative of the French metropolitan population. The size of the panel has significantly expanded since 2008: the panel was composed of about 13,000 households before 2008 but raised to more than 22,000 households after 2008.

KWP also provides a set of weights to correct for some bias of measure. A first one ensures the representativeness of the panel over time and corrects for some under-reporting behaviors within each month. A second one corrects for the quantities and expenditures recorded for purchases that benefit from indirect promotions (*e.g.*, discounts on loyalty cards). Finally, a third weight corrects for evident recording errors or exceptionally large quantity purchases. All these weights are used when computing the monthly average price of each chain-product pair.

B.2. Construction of the Brand Type Variable

The basic principle of our research design is to compare the changes in price of NB products and private labels products offered by traditional retailers, resulting from authorizing input price discrimination. It is therefore essential to classify products according to their brand type. The KWP data informs about whether a product is a private label (offered or not by a traditional retailer) or not. We then construct a procedure that enables us to classify each brand name into one of the following four brand types: national brands (NB), private labels (PL), discount private labels (PL-D), and first-price products (FP). This categorization follows the usual classification adopted in France.

Private labels are defined according to the Private Label Manufacturers' Association (PLMA) as "products [that] encompass all merchandise sold under a retailer's brand. That brand can be the retailer's own name or a name created exclusively for that retailer. In some cases, a retailer may belong to a group that owns the brands which are made available to all the members of the group".[54]

First-price products are varieties offered at the lowest price for a given product category. A first-price product can be either sold under a retailer's brand (denoted PL-FP in this particular case) or sold under a generic name. In the first case, the brand can be composed of the retailer's own name (*e.g.*, Carrefour Discount) or be created exclusively by that retailer (*e.g.*, Eco+ for Leclerc). It represents a low-cost version of the traditional retailer's private label. It is usually offered in the multiple chains of a retail group. In the last case, the FP product is produced by an independent manufacturer which choose a brand name. These products are usually sold in stores of competing retail groups.

The procedure presented below is defined at the brand name level (*e.g.*, Coca Cola Zero, Danone Activia, Heineken) which means that every products sold under the same brand name are classified in the same brand type. The procedure is composed of several steps that are applied in the following order:

1. A brand name whose sales are made in more than 90% of cases in a discount store is classified as a private label offered by discounter, *i.e.* a PL-D.

2. A brand name that is identified as a private label in the KWP data and that is not a PL-D is defined as a private label sold by a traditional retailer, *i.e.* a PL.

3. The brand name that satisfies at least one of the following criteria is identified as a brand name of first-price products,*i.e.* FP:

    - the brand name is composed of one of the two terms: "1er prix" or "premier prix" (French translation of "first price");

[54] https://www.plmainternational.com/industry-news/private-label-today.

- the brand name corresponds to one of the retailers' own brand dedicated to first-price products (*e.g.*, *Carrefour Discount* for Carrefour, *Eco+* for Leclerc, *Bien vu* for Système U);
- (i) the name of the brand is composed of "sans marque" (an equivalent for "no name") and (ii) the annual average price of the brand name is less than or equal to 95th percentile of the distribution of the annual average price of PL-D products in this category.[55]
- the annual average price of the brand name (not classified before) is less than or equal to the annual average price of PL-D and FP brands previously identified in this category.

4. Finally, the brand names that do not satisfy one of the previous criteria are considered as national brands, *i.e.* NB.

We only observe small variations over the period of study whatever the brand type. NB products have a market share of 65% on average. Private labels have on average a market share of almost 22%, 8% for private labels sold by discounters (PL-D) and 5% for FP products respectively. Summing the market shares of PL and PL-D, we obtain an average of 30% over the period.

B.3. Composition of the Final Sample and Descriptive Statistics

The list of the retail groups included in the final sample is reported in Table B.1 as well as their market share calculated from the total food expenditure recorded in the KWP and in the final sample.

We also provide some detailed summary statistics at the product category level in Table B.2.

Table B.1: List of Retail Groups and Market Shares

Retail group	Market Share	
	KWP	Final Sample
R1	16.37	20.01
R2	8.81	10.47
R3	13.48	15.81
R4	11.63	13.52
R5	25.05	24.46
R6	3.46	3.89
R7	12.57	10.98
R8	8.63	0.86

Notes: The table gives the average monthly market shares on food sales over the 2006-2010 period by retail group using both the KWP dataset and the final sample. The names of retail groups are not disclosed due to a confidentiality agreement with KWP.

[55] This corresponds to the average price plus two times the standard deviation.

Table B.2: Summary Statistics by Product Category

Product category	No. obs.	No. products	Average monthly price Min	Max	No. transactions	Total expenditure
Sweeteners	58327	293	0.21	83.35	343916	796036.56
Cider	15015	209	0.43	7.69	53254	158829.03
Soups	95489	828	0.08	16.48	341730	755007.38
Pickles	15374	145	1.24	21.25	87880	213269.28
Ready to eat desserts	10143	112	0.51	16.20	28632	94925.45
Cold sauces	92946	1024	0.23	86.61	373627	671820.75
Fruit syrup	38146	591	0.33	23.13	156628	375093.38
Frozen snails	1498	88	2.02	97.74	4271	40316.90
Frozen fruits, juice and puree	2389	79	1.14	35.41	6037	25749.45
Baby milk	20343	144	0.31	26.94	61607	908439.19
Baby food	30080	185	0.14	32.72	268902	652423.31
Season chocolate	37283	1351	0.13	190.78	169087	1098192.88
Fruits and salted nuts	70813	1198	0.78	108.41	275212	669521.88
Margarine	45216	174	1.02	18.94	386187	958805.00
Breakfast chocolate and tea	24855	159	1.56	35.26	137571	430551.53
Soft drinks	260240	2862	0.05	97.42	1587411	4478611.50
Cereal bars	23348	187	3.80	115.27	88027	219202.73
Chocolate bars	22763	160	1.20	94.95	110396	427933.22
Sweets	125304	1719	0.89	479.59	439371	1100105.75
Whipped cream	5253	33	2.50	18.14	45239	78959.02
Other frozen food	1541	53	0.77	27.05	4315	13786.34
Frozen soup	467	25	1.33	23.54	701	2660.13
Fruit in syrup	23549	350	0.42	35.23	77716	181561.11
Jelly	1382	21	2.87	566.14	4388	10096.09
Freeze-dried vegetables	2137	95	5.53	582.62	3728	15424.16
Frozen pastries	9269	513	0.14	145.39	18031	86120.91
Petits-suisses	29467	143	0.71	19.98	257098	643530.94
Dried meals	9023	160	0.70	59.22	22320	48690.42
Frozen meals and starters	52451	1110	0.96	91.75	183032	755663.06
Frozen fish, surimi and shellfish	40597	942	0.70	1878.23	148789	832308.50
Fish spread	23256	469	2.23	3151.15	81911	214886.05
Cooking mix for creams and mousses	13060	122	0.88	92.85	38125	78717.13
Cooking mix for desserts and cakes	19936	186	1.17	26.57	50591	161443.09
Tinned salads	4041	75	2.43	19.88	9476	21190.11
Hot sauces	83068	938	0.35	2458.49	331467	566034.25
Long-life yogurts and desserts	4801	48	0.57	10.15	19363	42687.41
Prepacked dry-cured ham	25581	491	2.70	112.27	112302	431174.41
Prepacked ham (w/o EAN)	9393	891	1.50	124.17	13404	43556.59
Cooked ham	28053	455	0.20	286.74	130260	475457.81
Champagne	9896	229	2.28	64.62	21400	763173.13

Continued on next page

Table B.2: Summary Statistics by Product Category (*Continued from previous page*)

Product category	No. obs.	No. products	Average monthly price Min	Average monthly price Max	No. transactions	Total expenditure
Other alcohols	25933	673	2.84	293.30	55057	532674.63
Whisky and whiskeys	25448	288	0.00	106.53	73236	1375249.13
Alcoholic drinks	47993	530	0.46	72.77	165801	1826842.38
Local wines	34697	979	0.06	32.84	98528	598673.00
Table wine and foreign wines	24055	452	0.19	27.62	73753	306047.59
Champagne (w/o EAN)	728	45	8.37	52.75	916	39682.75
Brie cheese	6473	76	2.36	24.06	47997	111416.55
Camembert cheese	32717	170	0.43	42.03	362164	784814.56
Cantal cheese	1205	22	5.97	17.23	4380	11007.00
Edam cheese	1637	31	3.88	15.81	7755	15902.49
Emmental, gruyere and appenzeller cheese	27663	235	2.21	62.25	510171	1317340.88
Hot cheese	5858	76	2.63	50.13	17799	68671.93
Melted cheese, fresh or salted	98313	480	1.45	53.36	612021	1370453.63
Gouda cheese	2396	52	1.52	32.78	13112	29138.13
Blue cheese	18513	144	4.27	44.62	149724	363489.91
Pont l'Eveque cheese	2940	32	4.50	26.95	7131	19666.82
Pies and puff-pastries to fill	4293	25	0.45	6.69	29896	74839.70
Confit	5922	88	3.05	66.19	28303	106487.81
Raw or cooked chicken pieces	17659	302	0.06	46.56	85859	387991.25
Breaded meat	37530	397	0.44	93.46	192229	605135.69
Fresh pies	16330	261	1.42	52.25	62498	223224.23
Plums	8545	166	1.59	129.57	37024	134885.00
Other poultry	1563	67	1.74	87.38	3840	32174.40
Whole turkey (w/o EAN)	621	44	1.90	99.83	950	14419.56
Raw piece of turkey (w/o EAN)	18519	403	0.73	57.36	68038	323532.66
Raw/cooked turkey pieces	7955	163	0.01	45.32	28631	147367.39
Cooked chicken	433	7	3.52	13.96	4202	30702.95
Raw chicken (whole)	975	27	1.71	11.83	4243	31204.24
Raw chicken (whole) (w/o EAN)	20645	523	0.05	54.41	58101	443843.03
Smoked chicken (whole) (w/o EAN)	1114	18	2.09	14.29	2196	12208.22
Roasted turkey	1877	35	1.96	25.57	5252	26654.20
Duck (pieces) (w/o EAN)	13613	275	1.17	69.89	34269	264076.16
Rabbit (w/o EAN)	11705	176	0.21	32.21	28608	241494.88
Other poultry (w/o EAN)	971	32	1.43	33.30	1774	17056.72
Packed foie gras (EAN)	6349	461	5.10	428.55	17973	240372.73
Frozen meat (w/o EAN)	925	44	0.70	71.50	1704	14654.78

Notes: This table reports detailed information at the product category level for the categories retained in the final sample. The statistics reported are: the number of chain-product-month triplets (No. obs.), the number of products (No. products), the minimum and maximum average monthly price of a chain-product pair, the number of purchase transactions (No. transactions), and total expenditures (Total expenditure). w/o EAN stands for without a European Article Number.

Online Appendix for "The Effect of Input Price Discrimination on Retail Prices: Theory and Evidence from France"

Marie-Laure Allain, Claire Chambolle & Stéphane Turolla

A. Alternative Starting Date of the LME Effect

We test whether our results are sensitive to the date on which the effect of input price discrimination is supposed to materialize. In the baseline specification, we have assumed that manufacturers have begun to price discriminate in the days following the adoption of the law. As discussed before, it is possible that these negotiations only concern a small part of the business and that the bulk of the effect occurs during the following round of negotiations. We therefore test whether our results are robust when moving forward the starting date of the LME.

It is worth mentioning that the choice of a different starting date entails some changes in the composition of the final sample. Moving forward the starting date generates a transitory period. Consequently, we remove from the final sample the observations between August 2008 and the assumed starting date, in order to leave the pre-LME period uncontaminated from a potential effect of the law. This entails a change in the composition of the final sample as we keep only the chain-product pairs that satisfy our selection criteria for the new period of study (see Section 5.1.3).

Table A.3 reports the results of the estimation of the baseline regression model using equation 12 and two different starting dates. As before, the first row corresponds to the baseline estimate for ease of comparison. In the second row, we move the starting date of the LME to November 2008, which corresponds to the first month of the 2008-09 annual round of negotiations. We find that shifting forward the date of the LME by 3 months does not alter the sign and the statistical significance of the price effect of authorizing input price discrimination. Further, the magnitude of the causal effect is very similar to the one obtained in the baseline scenario. In the third row, we adopt a more conservative approach and consider that the effect of input price discrimination does not materialize before the end of the 2008-09 negotiations, *i.e.*, March 2009.[56] Despite a transitory period of 8 months, we obtain a highly significant point estimate whose value is close to that of the baseline scenario.

[56]Note that this is a rather extreme scenario as usually retailers and suppliers agree to end the negotiation on the the terms of sale at the end of the civil year as it was the case in 2006 or 2007.

Table A.3: Alternative Time Frames

Period of study	Transitory period	Starting date	$\widehat{\beta}$ Coef.	S. E.	Obs.	R^2
Panel A: Baseline						
2006-2010	No	August 2008	-0.0262	0.0052	1,919,872	0.9893
Panel B: Starting date & transitory period						
2006-2010	2008/08 to 2008/10	November 2008	-0.0273	0.0055	1,780,214	0.9892
2006-2010	2008/08 to 2009/02	March 2009	-0.0257	0.0057	1,521,398	0.9893

Notes: Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The first row reports the point estimate obtained in the baseline scenario for ease of comparison. The change of the date on which the effect of input price discrimination is supposed to materialize requires to select a new sample of products for each sensitivity analysis. Standard errors (denoted S. E.) are clustered at the chain level.

B. Robustness Tests on the Trimming Procedure

We present in this section some robustness tests to assess the sensitivity of our results to the trimming procedure applied in the paper. The aim of the trimming procedure is to remove purchase transactions with abnormal unit prices which potentially indicates some measurement errors. Recall that in the baseline procedure, we exclude purchase transactions whose normalized deviation between the unit price and the monthly average price of the product (excluding the unit price considered) is greater than 10 in absolute value as well as purchase transactions whose unit price deviates from a similar magnitude from the average price of the chain-product pair.

As a first robustness test, we present in Panel B of Table B.4 the estimation results of the DiD regressions when no trimming procedure is applied. As before, Panel A reports the baseline estimates for ease of comparison. The comparison of the number of purchase transactions between Panel A and Panel B indicates that the trimming procedure identifies only 6,266 unit prices as abnormal. Once applying the selection criteria, we obtain a final sample composed of 1,185 (=1,921,091 - 1,919,906; see column 1) additional monthly chain-product pair observations. As a result, the estimates reported in Panel B are almost identical to those reported in Panel A.

In a second step, we conduct a series of robustness tests in which we vary the threshold above which unit prices are deemed abnormal. Panel C presents the estimation results where we exclude transactions whose unit price is among the 0.1% smallest and the 99.9% largest prices recorded for a product in a given month and for a product within a chain. In panel D, we enlarge the range of extreme values by removing prices in the 2.5 and 97.5 percentiles. Finally, in Panel E, we use an alternative trimming procedure based on the construction of box plots and we defined as outliers prices that are below (above) the first (third) quartile minus (plus) three times the interquartile range.

Table B.4: Robustness Tests on the Trimming Procedure

Dependent variable	(log) price (P_{ikt})		
		With monthly trend	
	Baseline	Chain	Category
	(1)	(2)	(3)
Panel A: Baseline Trimming Procedure			
Treatment × PostLME	-0.0160	-0.0162	-0.0262
	(0.0045)	(0.0047)	(0.0052)
R^2	0.9886	0.9890	0.9893
Observations	1,919,906	1,919,559	1,919,872
# Purchase transactions	9,388,537		
Panel B: No trimming procedure			
Treatment × PostLME	-0.0153	-0.0155	-0.0262
	(0.0045)	(0.0047)	(0.0052)
R^2	0.9884	0.9888	0.9891
Observations	1,921,091	1,920,776	1,921,084
# Purchase transactions	9,394,803		
Panel C: Outliers outside [0.1%; 99.9%]			
Treatment × PostLME	-0.0154	-0.0157	-0.0272
	(0.0044)	(0.0046)	(0.0052)
R^2	0.9884	0.9888	0.9892
Observations	1,904,672	1,904,363	1,904,665
# Purchase transactions	9,334,371		
Panel D: Outliers outside [2.5%; 97.5%]			
Treatment × PostLME	-0.0153	-0.0155	-0.0273
	(0.0045)	(0.0047)	(0.0052)
R^2	0.9885	0.9889	0.9892
Observations	1,888,041	1,887,729	1,888,033
# Purchase transactions	9,285,301		
Panel E: Outliers and Box Plots			
Treatment × PostLME	-0.0113	-0.0110	-0.0247
	(0.0046)	(0.0048)	(0.0057)
R^2	0.9904	0.9907	0.9910
Observations	1,840,761	1,840,408	1,840,753
# Purchase transactions	8,701,844		

Notes: The table presents the DiD estimation results obtained from final samples that differ in their construction due to alternative trimming procedures. As before, observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The sets of fixed-effects used in each regression are identical to those reported in Table 4 of the manuscript. Panel A gives the point estimates obtained in baseline regressions for ease of comparison. Standard errors (denoted S. E.) are clustered at the chain level.

Overall, we observe that removing a larger number of purchase transactions affects less than proportionally the number of chain-product-month triplets retained in the final sample. This indicates that the issue of abnormal prices principally concerns chain-product-month triplets with few purchase transactions.

More importantly, we find that our results are barely affected by removing a larger number of unit prices in the distribution tails. Using our preferred specification (see column 3), we obtain some point estimates that are not significantly different from the one obtained with the baseline procedure. We are thus confident that our results are robust to changes in the trimming procedure applied.

C. Parallel Trend Assumption: Additional Results and Robustness Tests

This section discusses the baseline selection procedure adopted to satisfy the parallel trend assumption within product category. Precisely, we test in our baseline procedure, for each product category, the existence of a specific linear trend of national brand prices compared to private label prices in the pre-LME period (see equation 10). In case of the rejection of the null hypothesis at the 5% significance level- the coefficient associated to the slope of the trend is equal to zero -, the product category is ruled out from the final sample. In subsection C.1, we show the robustness of our DiD estimates for alternative significance levels.

One potential shortcoming of this procedure is that it indirectly assumes that the prices of private labels and national brands follow a linear trend. However, if private label prices follow a non-linear trend, the conclusion that national brand prices do not have a specific linear trend is not sufficient to ensure that both types of products follow a similar price trend.

We first show in subsection C.2 that this concern might be limited. Indeed, we show graphically that the categories that have the largest sales have experienced linear trends parallel in the pre-LME period. In subsection C.3, we apply alternative selection procedures that directly address the issue of potential non linear price trends and confirm the robustness of our DiD results.

C.1. Alternative Significance Levels of the NB trend Coefficient

We run the same selection procedure using more or less stringent statistical significance thresholds: we exclude from the final sample all categories whose the p-value of the coefficient of the national brand price trend is below 10% or below 1%. The 10% threshold is a more conservative scenario compared to the baseline threshold of 5%, as we are able to reject the null hypothesis of the absence of a trend at a higher level. With the 10% threshold, we only retained 64 product categories in the final sample compared to 76 with the baseline threshold. The 1% threshold is less stringent and leads us to select 94 product categories in the final sample.

We report the estimations results of DiD regressions of these two alternative cases in Table C.5. Despite the changes in the composition of the final sample, we still observe

Table C.5: Parallel Trend Assumption and Alternative Significance Levels

Dependent variable	(log) price (P_{ikt})		
		With monthly trend	
	Baseline	Chain	Category
	(1)	(2)	(3)
Panel A: Baseline (threshold 5%)			
Treatment × PostLME	-0.0160	-0.0162	-0.0262
	(0.0045)	(0.0047)	(0.0052)
R^2	0.9886	0.9890	0.9893
Observations	1,919,906	1,919,559	1,919,872
Panel B: threshold 1%			
Treatment × PostLME	-0.0171	-0.0168	-0.0275
	(0.0046)	(0.0048)	(0.0053)
R^2	0.9878	0.9882	0.9886
Observations	2,488,767	2,488,423	2,488,730
Panel C: threshold 10%			
Treatment × PostLME	-0.0151	-0.0151	-0.0174
	(0.0040)	(0.0041)	(0.0042)
R^2	0.9893	0.9896	0.9898
Observations	1,739,800	1,739,445	1,739,766

Notes: Observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The sets of fixed-effects used in each regression are identical to those reported in Table 4 of the manuscript. Panel A gives the point estimates obtained in baseline regressions for ease of comparison. Standard errors (denoted S. E.) are clustered at the chain level.

that authorizing input price discrimination has reduced the price of national brands as compared to private labels regardless of the significance threshold retained.

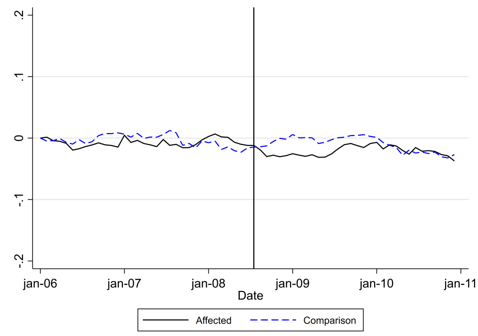
C.2. Some Additional Evidence on the Parallel Trend Assumption

We provide some additional information and graphical evidence regarding the parallel trend assumption in the pre-LME period for the product categories with the largest sales in the final sample. For that, we compare the monthly average variations of prices of the national brands and private labels for a given product category. These monthly variations are obtained by regressing the weighted-average of (log) prices on a set of monthly dummy variables for each brand type separately:

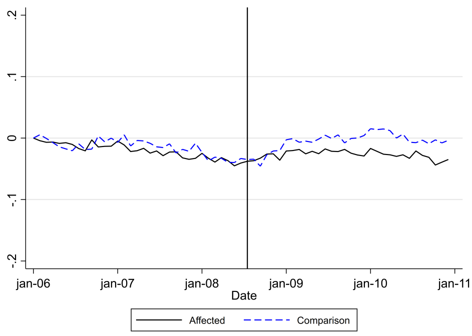
$$ln\left(P_{kit}\right) = \alpha + \sum_{l=2}^{59} \varphi_l Month_t^l + \mu_{ki} + \varepsilon_{kit}$$

where $Month_t^l$ are time period dummies and μ_{ki} are product-chain fixed effects. Observations are weighted by the expenditure shares of products in the total expenditure over the pre-LME period. The estimated coefficients $\widehat{\varphi}_l$ correspond to the difference in the average (log) prices between month l and the omitted first month of the period ($l = 1$). We then plot in Figures A1 and A2 the monthly price variations $\widehat{\varphi}_l$ of the national brands and private labels against the time period for the selected product categories.

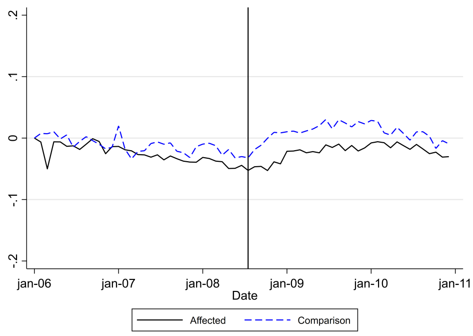
Figure A1: Price Trends for the Most Purchased Categories (Rank 1 to 6)



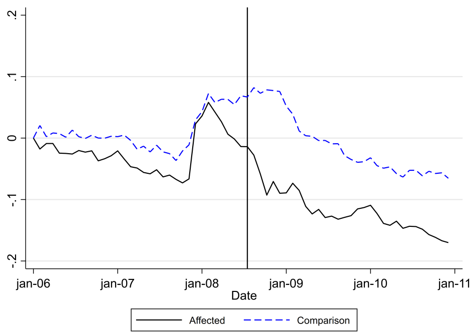
(a) Soft drinks



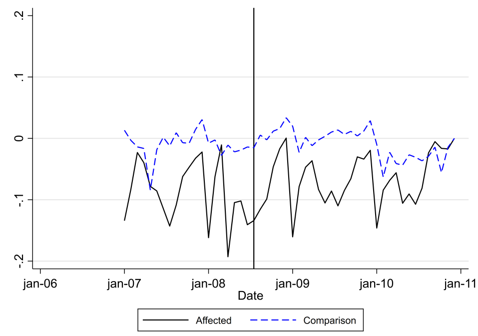
(b) Aperitif drinks



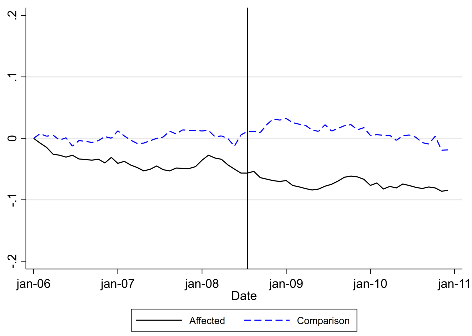
(c) Whiskey and bourbon



(d) Emmental, Gruyère, Appenzel



(e) Celebration chocolates

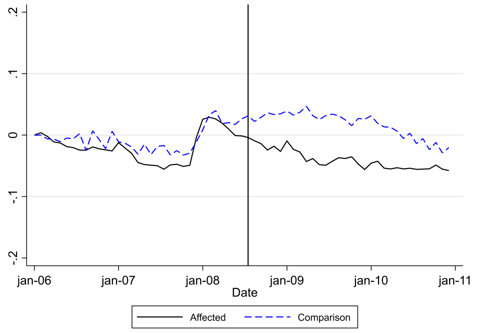


(f) Candies

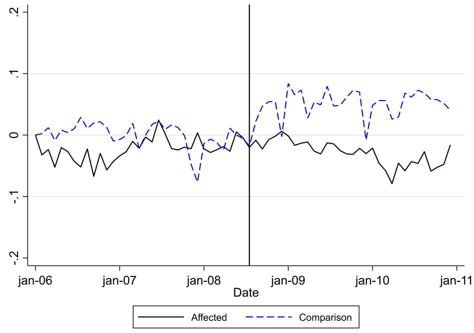
Notes: The figure presents the monthly average price variations of national brands and private labels, over the 2006-2010 period, for several product categories. The selected product categories belong to the most purchased categories over the pre-LME period. They are presented in descending order.

The figures show that the average prices of national brands and private labels have experienced parallel trends during the pre-LME period within a given product category, even if some monthly deviations between national brands and private labels are observable for

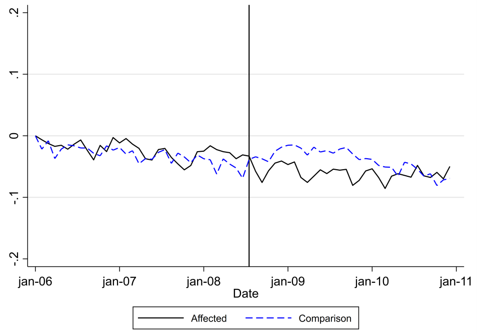
Figure A2: Price Trends for the Most Purchased Categories (Rank 7 to 10)



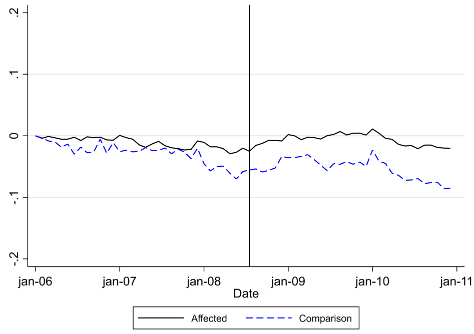
(a) Processed, fresh, salted cheeses



(b) Frozen fish, crustaceans, surimis



(c) Frozen dishes and starters



(d) Sweeteners

Notes: The figure presents the monthly average price variations of national brands and private labels, over the 2006-2010 period, for several product categories. The selected product categories belong to the most purchased categories over the pre-LME period. They are presented in descending order.

some categories - partly due to a composition effect generated by chain-product pairs that are not present every month.

The representation of the monthly price variations at the category level also shows that each category encounters a specific evolution, which stresses the importance of including category-month fixed effects in the DiD regressions. Overall, the graphs highlight that, after the LME reform, national brand prices have either experienced a larger drop or a lower rise than private labels.

As our analysis is mostly performed at the aggregate level, we now provide additional evidence that the parallel trend assumption is also satisfied at the aggregate level. We thus estimate equation (10) using all the data and test the statistical significance of the slope of a specific trend for national brands in the pre-LME period. The point estimate is not significant ($\widehat{\beta} = 0.0001(0.0002)$) which indicates that there is no significant difference between the trend of national brands and private labels in the pre-LME period.

C.3. Alternative Selection Procedures of the Product Categories

Linear Trend of the Monthly Deviations. We first apply the method proposed by Aguzzoni *et al.* (2016) to test the parallel trend assumption for each product category. The main advantage of this test is that it does not assume that the trend of all types of products' prices follows a linear trend in the pre-period. This test is composed of two stages.

First, we estimate the specific monthly deviations of national brand prices (relative to private labels) in the pre-LME period by replacing the continuous time variable by monthly dummies. This corresponds to the following OLS regression model:

$$\ln(P_{ijt}) = \alpha + \sum_{l=2}^{33} \varphi_l Month_t^l + \sum_{l=2}^{33} \gamma_l Month_t^l \times T_{ij} + \delta T_{ij} + \mu_{ij} + \varepsilon_{ijt} \tag{A-6}$$

where φ_l captures the average price deviation of private labels in month l compared to the initial month ($l = 1$). The coefficients γ_l measure the average price deviation of the affected chain-product pairs from the average price of the comparison chain-product pairs in each month.

Next, we estimate the slope of a linear trend of theses estimated monthly deviations. The statistical significance of the coefficient of the slope informs about a difference in the trends of national brand and private labels during the pre-LME period. It means that the average prices of national brands does not evolve around that of private labels, but significantly diverges - up or down - during a long spell of the pre-period.

We use this formal test to check whether the parallel trend assumption is satisfied at the product category level and to select the product categories which define a new sample. Panels B and C of Table C.6 present the results of the estimation of the DiD regressions when using a 1% or 5% significance threshold, respectively. These thresholds correspond to the significance level below which we reject the null hypothesis of the absence of a specific trend for national brands. With a less stringent significance threshold of 1%, we obtain a final sample composed of 100 product categories (compared to 76 in the baseline sample) with new product categories, some of them with high market sales, such as Beer, Milk and Coffee. This alternative sample covers 23.76% of the expenditures reported in KWP. With a 5% significance threshold we retain 83 product categories, which is still larger than the baseline sample.

Despite these composition changes in the sample, it is reassuring to observe that the relative price of national brands has significantly decreased after authorizing input price discrimination. The point estimates are still highly significant and close to the baseline values. This suggests that our baseline results are not driven by the potential selection

Table C.6: Alternative Selection Procedures of the Product Categories

Dependent variable	(log) price (P_{ikt})		
		With monthly trend	
	Baseline	Chain	Category
	(1)	(2)	(3)
Panel A: Baseline procedure			
Treatment × PostLME	-0.0160	-0.0162	-0.0262
	(0.0045)	(0.0047)	(0.0052)
R^2	0.9886	0.9890	0.9893
Observations	1,919,906	1,919,559	1,919,872
% of KWP expenditures	19.62		
Panel B: Slope of the monthly deviations (1% significance level)			
Treatment × PostLME	-0.0240	-0.0241	-0.0191
	(0.0037)	(0.0038)	(0.0038)
R^2	0.9900	0.9903	0.9906
Observations	2,654,258	2,653,961	2,654,244
% of KWP expenditures	27.35		
Panel C: Slope of the monthly deviations (5% significance level)			
Treatment × PostLME	-0.0229	-0.0232	-0.0196
	(0.0036)	(0.0037)	(0.0037)
R^2	0.9903	0.9906	0.9909
Observations	2,332,973	2,332,658	2,332,960
% of KWP expenditures	25.12		
Panel D: F-test on the monthly deviations (1% significance level)			
Treatment × PostLME	-0.0095	-0.0108	-0.0137
	(0.0056)	(0.0059)	(0.0054)
R^2	0.9757	0.9773	0.9764
Observations	622,756	622,251	622,750
% of KWP expenditures	7.71		
Panel E: F-test on the monthly deviations (5% significance level)			
Treatment × PostLME	-0.0105	-0.0141	-0.0151
	(0.0069)	(0.0074)	(0.0071)
R^2	0.9629	0.9657	0.9641
Observations	331,885	331,317	331,879
% of KWP expenditures	4.08		
Panel F: Visual inspection			
Treatment × PostLME	-0.0204	-0.0197	-0.0261
	(0.0044)	(0.0046)	(0.0046)
R^2	0.9896	0.9900	0.9905
Observations	4,073,613	4,073,327	4,073,607
% of KWP expenditures	39.40		

Notes: This table presents DiD estimation results using alternative selection procedures of the final sample. The selection procedure is run at the product category level and ensures that prices of national brands and private labels follow a parallel trend in the pre-LME period. Each panel contains estimations results obtained from a particular sample of products. As before, observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The sets of fixed-effects used in each regression are identical to those reported in Table 4 of the manuscript. Panel A gives the point estimates obtained in baseline regressions for ease of comparison. Standard errors (denoted S. E.) are clustered at the chain level.

of some product categories with a differential (non-linear) trend between national brand and private label prices.

This alternative selection procedure has the advantage to retain in the sample product categories for which the prices of national brands and private labels follow the same trends (whether linear or not). To test the parallel trend assumption at the aggregate level, we also apply the same test when aggregating all the products categories selected with this alternative selection procedure. For panel C, we fail to reject the null hypothesis of the absence of a specific trend for national brands at the 10% significance level. For panel B, we reject the null hypothesis of the absence of a specific trend for national brands only at the 10% significance level.[57]

Joint Significance of the Monthly Deviations. Another possible selection procedure follows the formal test of the parallel trends assumption proposed by Ashenfelter *et al.* (2013). It consists in testing the joint significance of the specific monthly price variations of the affected products. A failure to reject the null hypothesis provides support to the common trend hypothesis between affected and comparison products.

We apply this test at the product category level and retain in the final sample the product categories for which we fail to reject the null hypothesis of the joint nullity of the estimated coefficient of national brands' monthly deviations (*i.e.*, $\hat{\gamma}_2 = \hat{\gamma}_3 = \ldots = \hat{\gamma}_{33} = 0$). The final sample is composed of 38 or 24 product categories according to the chosen significance level (1% or 5%, respectively). This corresponds to a sharp reduction of the sample size compared to the baseline sample (76 categories). This statistical test thus seems to impose too much constraints to satisfy the parallel trend assumption. Indeed, with a large number of months in the pre-LME period (*i.e.*, 32 without the reference month), this test unreasonably rejects the null hypothesis of a common trend.

Despite the stringency of this selection procedure, it is reassuring to observe that our results hold. Panels D and E of Table C.6 present the DiD estimation results. For the two final samples considered, we find that authorizing input price discrimination has significantly, though less significantly, reduced the relative price of national brands (in an order of -1.4% and -1.5%).

Visual Inspection. In a last robustness test, we follow the suggestions of Angrist and Pischke (2008) and select the product categories included in the final sample on the basis of a visual inspection of their price trend of national brands and private labels. In doing so we obtain a final sample composed of 80 product categories but above all of more than

[57]These approaches offer interesting alternative selection procedures. However, we prefer to keep our baseline procedure because when using our preferred specification with category-month fixed effects the parallel trend assumption is not well satisfied during the first year of the pre-LME period.

4 millions of chain-product-month observations that represent a total of 39.40% of KWP expenditures (see Table C.7 for the list of product categories).

The DiD estimation results are reported in Panel F of Table C.6 and again confirm the robustness of our results. While this selection procedure seems interesting to use when price trends follows complicated patterns, again it does not ensure the validity of the parallel trend assumption at the aggregate level. Checking that national brands do not experience a specific linear trend in the pre-LME period, we find that the parallel trends assumption is rejected at a 1% significance level.

Table C.7: Visual Inspection and List of the Product Categories Retained

Sweeteners	Petits-suisses
Cider	Dried meals
Rice	Frozen meals and starters
Soups	Fish spread
Fruit puree	Cooking mix for desserts and cakes
Pickles	Hot sauces
Beer	Meal substitutes
Vinegar	Yogurts
Stocks and cooking aids	Long-life yogurts and desserts
Ready to eat desserts	Prepacked dry-cured ham
Milk to drink	Prepacked ham
Chocolate bars	Fresh pizza
Semolina and polenta	Cooked ham
Cold sauces	Raw knuckle of ham (w/o EAN)
Coffee and chicory	Champagne
Baby food	Punchs
Jam	Cocktails
Baby flour	Other alcohols
Fruits and salted nuts	Whisky and whiskeys
Fat	Alcoholic drinks
Margarine	Local wines
Dry bread	Table wine and foreign wines
Dry pasta	Fresh deli
Breakfast chocolates and tea	Other cheese
Baby breakfast	Brie cheese
Soft drinks	Goat cheese
Butter	Melted cheese, fresh or salted
Whipped cream	Munster cheese
Crème fraiche	Blue cheese
Frozen cooked vegetables	Pont l'Eveque cheese
Frozen raw vegetables and herbs	Confit
Frozen french fries and potato pancakes	Raw or cooked chicken pieces
Other frozen food	Raw sliced bacon and pork belly
Fresh pasta	Breaded meat
Canned pate and rillettes	Fresh quenelles
Olives	Raw piece of turkey (w/o EAN)
Appetizer	Raw/cooked turkey pieces
Cookies	Raw chicken (whole) (w/o EAN)
Breakfast cereals	Honey
Fruits in syrup	Industrial pastry

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