The Effect of Input Price Discrimination on Retail Prices: Theory and Evidence from France*

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Abstract

A ban on input price discrimination directly impacts suppliers who sell through competing retailers (e.g. national brand suppliers) but only indirectly those who sell through one exclusive retailer (e.g. private label suppliers). In a secret contracting environment, we show that, because of opportunism, removing a ban on input price discrimination reduces the retail price of the national brands. In contrast, removing the ban on input prices has an ambiguous impact, though more limited, on the retail prices of private labels. A reform authorizing input price discrimination took place in France in 2008 and our paper uses this natural experiment to test our predictions. Using a consumer panel dataset of food prices in France over the period 2006-2010, we run a difference-in-differences analysis and show that on average the reform has led to a decrease in prices of national brands by 3.36% compared to private labels.

Keywords: Input Price Discrimination, Policy Evaluation, Food Retail Sector.

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

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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 its buyers when the buyers themselves compete to resell to consumers. In practice, input price discrimination arises because it is always 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 welfare effect of input price discrimination are likely to be ambiguous. However, an additional effect arises in the case of input price discrimination: buyers with high bargaining power may force upstream suppliers to offer advantageous wholesale price (typically the case in the food retail sector) which may in turn harm retail competition and consumer surplus.

Competition authorities have adopted various legislations to address their concern about input price discrimination. In the U.S., 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". Its intent was to preserve downstream competition by protecting small against large buyers. In Europe, Article 102 of the Treaty on the Functioning of the European Union prohibits a dominant firm from "applying dissimilar conditions to equivalent transactions with other trading parties, thereby placing them at a competitive disadvantage". Going one step further, France adopted in 1986 a specific regulation that forbids any supplier to offer different conditions to similar buyers. In 2008, the "Loi de Modernisation Economique" lifted this constraint in the case of the relationship between suppliers and retailers. The non-discrimination principle was suppressed in the retail sector with the intent to intensify competition among buyers and to reduce consumer prices. Finally, in 2018, the Norwegian competition authority launched an investigation on the relevance of instoring a ban on input price discrimination, after an internet shopping platform, that had recently entered, complained he could not compete on equal terms with incumbent retailers that were benefiting from advantageous input prices.[1]

The present paper takes advantage of the 2008 change in input price discrimina-

[1] See the press release at https://konkurransetilsynet.no/investigating-competition-issues-in-the-norwegian-grocery-sector/?lang=en

tion regulation in France to empirically assess its effect on prices. As a large strand of the theoretical literature predicts that the effect of intermediate price discrimination on prices can go either ways, an empirical approach is particularly helpful to determine which force is likely to dominate. Our paper attempts to do so and addresses several challenges. First, the change in regulation that we analyse was introduced at the national level and directly applied to all producers and retailers. Such a natural experiment does not provide a perfect control group that would enable a simple estimation of the causal effect of the reform on prices. We are therefore building an original theoretical model that allows us to form testable predictions. Second, we take advantage of a large consumer panel dataset (Kantar *TNS Worldpanel*) recording, at the household product level, all consumer food purchases and prices at the stores. We are therefore able to include a large range of food products in our empirical analysis using a robust difference-in-differences method.

Our original model is designed to represent a standard producer-retailer relationship. At the upstream level, we consider a national brand producer and two private label producers. At the downstream level, there are two imperfectly competing downstream firms. The national brand seller offers its good to the two competing retailers whereas each private label producer has an exclusive dealing relationship with one retailer. This model enables us to distinguish between a producer that is directly affected by the reform, the national brand seller, and the private label producers who cannot price discriminate because of the exclusive dealing assumption, and therefore are not directly affected by the law. We show that, with non linear secret contracts, authorizing input price discriminationleads to a decrease in the price of the national brand, and has an indirect, thus more limited, effect on the price of private label products. We then build our empirical approach on this result by considering that private labels, provided that they are indeed offered by a dedicated manufacturer, constitute a good candidate as a comparison group. We develop a difference-in-differences analysis over the period 2006-2010 around the change in regulation that took place in january 2009. We highlight that suppressing the ban on input price discrimination indeed lowered national brand food prices by 3.36% on average compared to the private labels.

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 enables us to form testable predictions about the effect of input price discrimination on prices. Section 4 gives some background on the French food retail sector and relevant legislation. Section 5 presents the data and puts forward the empirical strategy. Section 6 derives our empirical results, while robustness checks are provided in Section 7. Section 8 concludes.

2 Related literature

The theoretical literature on price discrimination is dense and brings out contrasted results. The assumptions on tariffs and contracts' observability play a key role and we present here the main results of the literature along these lines.

□ Public contracts – In vertically related market with public linear wholesale contracts, a "standard view" is that banning input price discrimination improves allocative efficiency and welfare. DeGraba (1990) show this in a framework with a monopolist supplying two Cournot competing downstream firms with asymmetric retail marginal costs. The supplier offers take-it-or-leave-it linear tariffs to downstream retailers and consumers' demand is linear. If the monopolist can price discriminate, the unit wholesale price offered in equilibrium to the efficient downstream firm is higher than that of its less efficient rival, because the firm with a lower marginal cost has the more inelastic demand for the input. Discrimination thus generates inefficiency.[2] However, the economic literature has offered a lot of contradicting arguments to this "standard view". For instance, Katz (1987) shows that discrimination may increase welfare when downstream firms can, at a fixed cost, integrate backward and supply at a given marginal cost (higher than the monopolist marginal cost). Because of the fixed cost, it is less costly for the more efficient firm who sells more units to integrate backward, and this threat enables it to obtain the lower wholesale unit price. In that case, discrimination, by avoiding inefficient backward integration, improves welfare.[3] Arya and Mittendorf (2010) analyze the effects of a ban on input price discrimination across retailers in a set-

[2]In the long run, DeGraba (1990) allows firms to invest to lower their level of marginal cost and shows that the discrimination now harms welfare.

[3]Inderst and Valletti (2009) revisit both DeGraba (1990) and Katz (1987). They extend Katz (1987) by considering that it is viable for the two downstream firms to integrate backward. They also follow the results of DeGraba (1990) by considering both short run and long run implications of a ban on input price discrimination. They confirm the result of DeGraba (1990) i.e. the ban is welfare improving in the short run but not in the long run.

ting where retailers are asymmetric and one operates in multiple markets. They find that price discrimination leads to price cuts in markets with lower demand and that, when these low demand markets are also less competitive, price discrimination can provide welfare gains by increasing the output on these markets. Inderst and Shaffer (2009) show that with observable two-part tariffs, the monopolist has an incentive to offer a lower wholesale unit price to the more efficient downstream firm because this is efficient for total industry profit. Under a ban, the monopolist raises both wholesale prices and relatively more that of the more efficient firm, and these two effects reduce welfare. Miklos-Thal and Shaffer (2018) focus on price discrimination across markets rather than across buyers, and show that discrimination may have a positive allocation effect: welfare can rise even if total output decreases.

Recent arguments in the literature have also conforted the "standard view'".Herweg and Muller (2014) highlight that when retailers are privately informed about their efficiency and that a manufacturer offers a menu of non linear contracts as a screening device, non discriminatory laws may improve welfare. A recent literature on platforms and "price parity clause" also offers a new angle of analysis on input price discrimination. Johansen and Vergé (2017) analyse the effect of such price parity clause in the framework where competing sellers distribute their products directly as well as through competing intermediation platforms (such as the hotel industry). They show that because of the sellers' participation constraints, price parity clauses may simultaneously benefit all the actors (platforms, sellers and consumers), even in the absence of traditional efficiency arguments.

□ Secret contracts – According to Rey and Tirole (2007), when contracts are secrets, opportunism may prevent a monopolist who sells to competing downstream firms to obtain the monopoly profit: each retailer fears that the monopolist grants a discount to its rival and therefore, only accepts a unit input price at marginal cost. This opportunism problem arises both if the monopolist offers take-it-or-leave-it unit wholesale prices or non linear contracts. Banning discrimination may facilitate the exercise of monopoly power because it enables the monopolist to commit not to offering secret price cut to a retailer's rival, and therefore equilibrium input and retail prices increase. O'Brien and Shaffer (1994) and O'Brien (2014) confirm that a ban on input price discrimination would lower welfare in a bargaining setting. The literature brings less contrasted results under secret contracts. An exception is Caprice (2006), who shows that if the monopolist competes with a less efficient competitive fringe, the ban could instead increase welfare. Under the ban, it would then be optimal for the monopolist, in order to exclude the fringe, to commit on a unit input price below cost in exchange for a higher fixed fee, which would decrease final prices.

□ Empirical literature – Although the theoretical literature is abundant, there is, to our knowledge only few empirical literature on input price discrimination. Villas-Boas (2009) develops a model with public unit wholesale contracts and simulates the effect of banning price discrimination on the wholesale market for coffee in Germany. She highlights that such a legislation would have welfare improving effects on that market. Hastings (2009) uses a model to represent the vertical channel in the gazoline market and simulates equilibrium prices under price discrimination and uniform wholesale pricing. She finds that average prices would rise five cents per gallon under uniform wholesale pricing. Grennan (2013) develops a structural model of secret bargaining and shows that, according to the theoretical predictions, more uniform prices soften competition on the input markets among hospitals.

Our model originally departs from the existing literature both theoretically and empirically. Our theoretical model is in line with the literature on secret contracts but originally takes into account the coexistence of suppliers that are affected differently by input price discrimination rules. Indeed, a ban on input price discrimination directly impacts suppliers who sell through competing retailers (e.g. national brand suppliers) but only indirectly those who sell through one exclusive retailer (e.g. private label suppliers). We show that, because of opportunism, removing a ban on input price discrimination reduces the retail price of the national brands. In contrast, removing the ban on input prices has an ambiguous impact, though more limited, on the retail prices of private labels. Empirically, the few existing papers that have analysed this issue have all developed structural econometric approaches and we are the first to actually benefit from a true natural experiment with available data before and after which allows a retrospective analysis. Moreover, whereas existing studies were focusing on one specific product, our dataset that record all food purchases of a representative sample of households provides a unique opportunity to evaluate the price effect of input price discrimination at a larger scale.

3 The Impact of input price discrimination on Intermediate and Retail Prices

In a market with a monopolistic supplier who sells its product through differentiated retailers, O'Brien and Shaffer (1994) show that a ban on input price discrimination unambiguously leads to an increase in wholesale and retail prices. In their framework, the suppliers offers secret two-part tariff contracts to the retailers. Under discrimination, fear of opportunism (see Hart and Tirole, 1990) drives each retailer to reject any tariff with a wholesale 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 wholesale price offered to the rival, hence suppressing any scope for opportunism. As a result, under such a ban, wholesale prices increase up to the point where retail prices reach the monopoly level. [4]

In this section we build on this approach and extend the analysis of input price discrimination to a more complex framework with upstream competition. We model the relationships between a duopoly of differentiated retailers supplied by differentiated producers. Among those suppliers, one sells to several retailers, while the others sell exclusively to one retailer. The possibility to price-discriminate on the input market will thus directly affect only the wholesale prices of the producer who supplies several retailers; however, all retail prices may be indirectly affected. In this setup, we derive theoretical predictions on the impact of removing a ban on input price discrimination on the equilibrium outcomes.

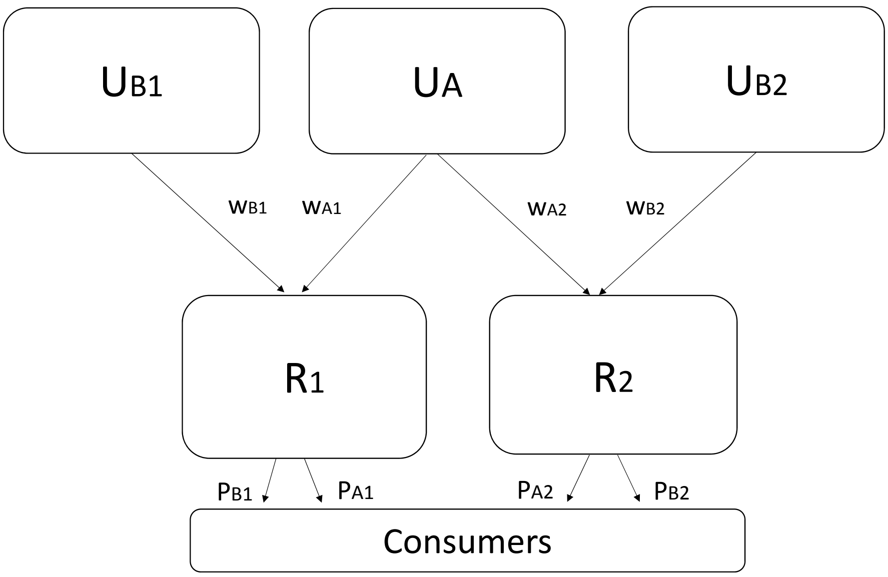
3.1 The model

Consider two differentiated retailers denoted R_i, with $i \in \{1, 2\}$ who can sell two differentiated products $k \in \{A, B\}$ to consumers. Good A is produced at a constant marginal cost c (with $0 \leq c$) by a manufacturer U_A who sells to the two retailers. By contrast, good B is produced by two independent suppliers U_{Bi}. Each supplier U_{Bi} sells the good exclusively to one retailer, R_i (see figure 1). Overall, due to retailers differen-

[4] This result is also implicitly shown by McAfee and Schwartz (1994) when they analyze the case of symmetric beliefs. They show that when each retailer believes his rival receives the same offer from the producer, opportunism is solved and the monopolist is able to capture its monopoly profit.

tiation, there are four products ki available to consumers. Product A may represent a national brand whereas each product Bi may represent the private label product sold by retailer R_i.

Figure 1: Market Structure



Products A and B are horizontally differentiated.[5] For the sake of simplicity, we assume that consumers' demand for product ki is symmetric accross retailers and across products.

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

We assume that:

Assumption 1 *Demands are downward sloping, and products are substitutes:*[6]

$$\frac{\partial D_{ki}}{\partial p_{ki}} < 0, \frac{\partial D_{ki}}{\partial p_{li}} > 0, \frac{\partial D_{ki}}{\partial p_{kj}} > 0$$

[5]In the specific case of private label, it is usual to assume vertical differentiation between the private label and national brand products. However, press release tend to show that private label products has increased over time. Furthermore, introducing vertical differentiation would make the equilibrium computation more tedious without affecting the main insights.

[6]Note that we do not make any *a priori* assumption on the sign of $\frac{\partial D_{ki}}{\partial p_{lj}}$.

For any given price vector $p = (p_{A1}, p_{B1}, p_{A2}, p_{B2})$,

a unit increase in both p_{ki} and p_{kj} causes the demand for ki to fall:

$$|\frac{\partial D_{ki}(p)}{\partial p_{ki}}| > |\frac{\partial D_{ki}(p)}{\partial p_{kj}}|$$

a unit increase in both p_{li} and p_{lj} causes the demand for ki to rise:

$$|\frac{\partial D_{ki}(p)}{\partial p_{li}}| > |\frac{\partial D_{ki}(p)}{\partial p_{lj}}|.$$

Let $D^3_{ki}(p_{ki}, \infty, p_{kj}, p_{lj})$ be the demand for product k at retailer R_i when it only sells product k. Assumption 2 also applies to $D^3_{ki}(p_{ki}, \infty, p_{kj}, p_{lj})$.

In some cases, we will make the following assumption, which guarantees that an increase in the price of product k sold by retailer R_i impacts more, on absolute terms, the total sales of product k than the total sales of the rival product l:

Assumption 2

$$|\frac{\partial D_{ki}}{\partial p_{ki}} + \frac{\partial D_{kj}}{\partial p_{ki}}| > |\frac{\partial D_{li}}{\partial p_{ki}} + \frac{\partial D_{lj}}{\partial p_{ki}}|$$

We consider the following two-stages game:

- Stage 1: Suppliers simultaneously offer take-it-or-leave-it secret contracts to their retailers. The wholesale contract offered by supplier U_k to retailer R_i consists of a unit price w_{ki} and a fixed fee F_{ki}. Each retailer then accepts or rejects the offer. Retailers have passive beliefs.[7] We also assume that 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. A retailer who rejects the offer cannot sell the good.

  When discrimination is prohibited, the national brand producer must offer the same unit input price to the two retailers: $w_{A1} = w_{A2} = w_A$.

- Stage 2: The two retailers compete by simultaneously setting their final prices p_{ki}. Wholesale contracts are not observed by the downstream competitor between the contract stage and the final price competition stage. By contrast, when discrimination is prohibited, the two retailers know that they have received the same unit input price. The unit input price for product A thus becomes observable.

[7] When a retailer receives an unexpected offer, he still believes that his competitor received the equilibrium offer.

First, note that our assumption that upstream firms offer take-it-or-leave-it contracts to retailers in Stage 1 is without loss of generality.[8] We adopt the contract equilibrium concept developed by Crémer and Riordan (1987). We could alternatively adopt a Nash-in-Nash bargaining process between upstream and downstream firms (with balanced bargaining powers), or even assume that retailers offer take-it-or-leave-it contract instead. All these settings are equivalent and would give the same result, i.e. the same effect of input price discrimination on wholesale and final prices.

We assume that discrimination over the fixed fees is not prohibited. In practice, it may be difficult for a court to establish that discriminatory fixed fees have been employed because such fees may retribute 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 case.

We now solve the game backward and assess the impact of a ban on input price discrimination on the equilibrium outcomes.

3.2 Price competition stage

We consider first the case in which input price discrimination is allowed. Assume that supplier U_A has offered different contracts (w_{Ai}, F_{Ai}) to the two retailers and that each supplier of private label product Bi offers his customer a contract (w_{Bi}, F_{Bi}).

We denote R_i's profit by :

$$\pi^i \equiv \sum_{k=A,B} (p_{ki} - w_{ki}) D_{ki}(p_{ki}, p_{li}, p_{kj}, p_{lj}) - F_{ki} \tag{1}$$

We denote by π^i_1 the derivative of retailer 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 notation extends to higher order derivatives.

We make the following regularity assumptions on the profit functions:

[8] Assuming in-house production by the retailer, or production by a competitive fringe, would not impact our results. The literature on private labels (cf. Bergès-Sennou, Bontems, and Réquillart (2004) for a survey), 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: these two assumptions lead to cost-based tariffs. Alternatively, private labels may also be produced by large manufacturers who produce both their national brand and private labels (See PLMA International, 2018, and Chambolle, Christin, and Meunier (2006)).

Assumption 3 *(i) For* $i = 1,2$*,* $\pi^i_{11} < 0$*,* $\pi^i_{22} < 0$*,* $|\pi^i_{11}| \geq \pi^i_{21} > 0$*, and* $|\pi^i_{22}| \geq \pi^i_{12} > 0$*.*
(ii) For $i = 1,2$ $\pi^i_{14} \leq 0$*,* $\pi^i_{13} \geq 0$ *with* $|\pi^i_{14}| < \pi^i_{13}$*,* $\pi^i_{24} \geq 0$*,* $\pi^i_{23} \leq 0$ *with* $|\pi^i_{23}| < \pi^i_{24}$*.*

Part (i) of Assumption 3 ensures that the retailer's profit functions are concave in prices and that the best response functions increase in the rival's prices (see Appendix A.1 for details). Part (i) of Assumption 3 guarantees the existence and uniqueness of the price equilibrium. Part (ii) is a regularity assumption that ensures that the marginal profit of a retailer on product k is positively (resp. negatively) affected by an increase in the price of product k (resp. product l) at its rival and that cross effects are smaller than direct effects.

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

$$\begin{aligned} \pi^i_1 &= D_{Ai}(.) + (p_{Ai} - w_{Ai})\frac{\partial D_{Ai}(.)}{\partial p_{Ai}} + (p_{Bi} - w_{Bi})\frac{\partial D_{Bi}(.)}{\partial p_{Ai}} &= 0 \\ \pi^i_2 &= D_{Bi}(.) + (p_{Ai} - w_{Ai})\frac{\partial D_{Ai}(.)}{\partial p_{Bi}} + (p_{Bi} - w_{Bi})\frac{\partial D_{Bi}(.)}{\partial p_{Bi}} &= 0 \end{aligned} \tag{2}$$

The two FOCs in system (2) define the best response functions denoted $p^r_{ki}(w_{ki}, w_{li}, p_{kj}, p_{lj})$. Their intersection is denoted $p^*_{ki}(w_{ki}, w_{li}, w_{kj}, w_{lj})$. Whether wholesale prices are observable or not $p^*_{ki}(.)$ defines the equilibrium final prices because when unobserved, the rival's input prices are consistently anticipated. As a result, if removing the ban on input price discrimination does not change the wholesale prices, it will not affect equilibrium retail prices.

In what follows, we make the following regularity assumption, which implies that a unit increase in the prices of product k at both retailers –due for instance to a cost shock on product k– affects more the marginal profit of a retailer on this product k than the marginal profit of the retailer on the rival product l:

Assumption 4

$$\begin{aligned} |\pi^i_{24} + \pi^i_{22}| &> \pi^i_{14} + \pi^i_{12} \\ |\pi^i_{13} + \pi^i_{11}| &> \pi^i_{23} + \pi^i_{21}. \end{aligned}$$

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

Lemma 1 *Under Assumption 1-4, we have* $\frac{dp^*_{Ai}}{dw_A} > 0$*, whereas, the sign of* $\frac{dp^*_{Bi}}{dw_A}$ *is ambiguous but* $|\frac{dp^*_{Bi}}{dw_A}| < |\frac{dp^*_{Ai}}{dw_A}|$*.*

Proof. See the Appendix A.2. ■

The effect of w_A on p_A is unambiguously positive which is intuitive. The effect of an increase in w_A on P_B is ambiguous. Considering the retail price decision of R_i by provide some insights. Given (p_{Aj}, p_{Bj}), in the plan (p_{Ai}, p_{Bi}), the optimal prices $p^r_{ki}(w_{ki}, w_{li}, p_{kj}, p_{lj})$ set by R_i are defined by the intersect between two increasing functions $p_{Ai}(p_{Bi})$ and $p_{Bi}(p_{Ai})$ implicitely defined by two equations of 2. An increase in w_A shifts upward $p_{Ai}(p_{Bi})$ but shifts downward $p_{Bi}(p_{Ai})$, though to a lesser extent under Assumption 1. As a result, the intersection of the two functions shifts to the right but can either be higher or lower than the initial point. An increase in w_A generates on the one hand the indirect effect through an increase in p_{Ai} that shifts P_{Bi} upward because interbrand competition is softened. On the other hand, for a given p_{Ai} the direct effect of an increase in w_A clearly shifts P_{Bi} downward because it decreases the margin on product A which in turn pushes the retailer to set a lower price for B in order to report demand from product A to product B on which the margin becomes relatively higher.

3.3 Contract stage under discrimination

We now look for the contract equilibrium when discrimination is allowed.

Let w^d denote the anticipated equilibrium vector of wholesale prices with discrimination $w^d = (w^d_{Ai}, w^d_{Bi}, w^d_{Aj}, w^d_{Bj})$.

Consider first the offer made by supplier U_{Bi} to retailer R_i. The two firms anticipate that R_j sets the equilibrium $p^*_{Bj} = p^*_{Bj}(w^d)$ and $p^*_{Aj} = p^*_{Aj}(w^d)$ defined by (2). Both U_{Bi} and retailer R_i anticipate the continuation equilibrium where R_i sets the prices $p^r_{Ai}(w^d_{Ai}, w_{Bi}, p^*_{Aj}, p^*_{Bj})$ and $p^r_{Bi}(w_{Bi}, w^d_{Ai}, p^*_{Aj}, p^*_{Bj})$.

U_{Bi}'s program is to maximize the following profit:

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

where $\overline{\pi}^i_A$ is the status quo profit of retailer R_i when he rejects producer U_{Bi}'s offer and accepts the equilibrium offer by supplier U_A. In the continuation equilibrium, R_i only sells product A at a price $p^{r3}_{Ai} \equiv p^r_{Ai}(w^d_{Ai}, +\infty, p^*_{Aj}, p^*_{Bj})$, and his profit $\overline{\pi}^i_A = (p_{Ai} - w^d_{Ai})D^3_{Ai}(p^{r3}_{Ai}, p^*_{Aj}, p^*_{Bj}) - F_{Ai}$ is independent of w_{Bi}.

Lemma 2 *When discrimination is allowed, under Assumptions 1-3, there is unique symmetric equilibrium $w^d_{Bi} = c$ for $i = 1, 2$.*

Proof. The FOC of supplier U_{Bi}'s programme (3) is:

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

Using (2) and simplifying yields:

$$0 = (w_{Bi} - c)\tfrac{\partial D_{Bi}}{\partial p_{Bi}} \tfrac{dp^r_{Bi}}{dw_{Bi}} + (w_{Bi} - c)\tfrac{\partial D_{Bi}}{\partial p_{Ai}} \tfrac{dp^r_{Ai}}{dw_{Bi}}.$$

Which cancels for $w_{Bi} = c$. The proof of unicity is provided in Appendix A.3. ■

Consider now the offer made by supplier U_A to retailer R_i. As U_A is assumed to send distinct delegates to both retailers, both U_A and retailer R_i anticipate the continuation equilibrium where R_i sets the prices $p^r_{ki}(w_{ki}, w^d_{li}, p^*_{Kj}, p^*_{lj})$ for $k = A, B$ and $l \neq k$. Hence, as U_B, U_A offers cost-based tariffs in equilibrium.[9]

Lemma 3 *When discrimination is allowed, under Assumptions 1-3 there is a unique symmetric equilibrium where $w^d_{A1} = w^d_{A2} = c$.*

Proof. See Appendix A.4. ■

Note that this equilibrium is unique whenever a retailer's best response to an increase in the unit wholesale price of product A is to decrease the price of the rival product (*i.e.* $\frac{dp^r_{Bi}}{dw_{Ai}} < 0$), which is true under Assumption 1-3. Therefore all wholesale prices being set to marginal cost, we obtain the following lemma:

Lemma 4 *When input price discrimination is allowed, under Assumptions 1-3, there is a unique symmetric retail price equilibrium $p^* \equiv p^*_{ki}(c, c, c, c)$ for $k = A, B$ and $i = 1, 2$.*

[9]This result is common in the literature on interlocking relationships with two-part tariffs. See for instance Allain and Chambolle (2011).

Proof. See Appendix A.4. ■

Again, this equilibrium is unique whenever a retailer's best response to an increase in the unit wholesale price of product A is to decrease the price of the rival's product (*i.e.* $\frac{dp^r_{Bi}}{dw_{Ai}} < 0$) which is true under Assumption 1-3.

3.4 Contract stage under a ban on discrimination

Consider the offer made by supplier U_{Bi} to retailer R_i. Here w^d_{Ai} is now replaced by w^{nd}_A. We denote the anticipated equilibrium wholesale price vector in that case $w^{nd} = (w^{nd}_A, w^{nd}_{Bi}, w^{nd}_{Bj})$. The pair $U_{Bi} - R_i$ now anticipates that R_i will adapt its prices according to $p^r_{Bi}(w_{Bi}, w^{nd}_A, p^*_{Bj}, p^*_{Aj})$ and $p^r_{Ai}(w^{nd}_A, w_{Bi}, p^*_{Aj}, p^*_{Bj})$. Supplier U_{Bi}'s program is the following:

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

The outside option profit $\overline{\pi}^i_A$ does not depend on w_{Bi}. Supplier U_B's program remains similar to (3) under the ban on discrimination. This leads to the following lemma:

Lemma 5 *When discrimination is banned, under Asssumptions 1-3 there is unique symmetric equilibrium $w^{nd}_{Bi} = c$ for $i = 1, 2$.*

Consider now the contracts offered by supplier U_A to both retailers. As mentioned earlier, under the ban the wholesale unit price received by R_i is the same as the one received by its rival, which implies that the negotiated wholesale contract between U_A and R_i is no longer secret. U_A, R_1 and R_2 anticipate the continuation equilibrium in which both retailers R_i and R_j adapt their prices to w_A, i.e. R_i ($i = 1, 2$) sets the prices $p^*_{Ai}(w_A) = p^*_{Ai}(w_A, w^{nd}_{Bi}, w_A, w^{nd}_{Bj})$ and $p^*_{Bi}(w_A) = p^*_{Bi}(w^{nd}_{Bi}, w_A, w^{nd}_{Bj}, w_A)$.

U_A's program is the following:

$$\begin{aligned} &\underset{w_A,F_{Ai}}{Max} \sum_i (w_A - c)D_{Ai}(p^*_{Ai}(w_A), p^*_{Bi}(w_A), p^*_{Aj}(w_A), p^*_{Bj}(w_A)) + F_{Ai} \\ &s.t., \texttt{for } i = 1, 2, F_{Ai} \leq (p^*_{Ai}(w_A) - w_A)D_{Ai}(.) + (p^*_{Bi}(w_A) - w^{nd}_{Bi})D_{Bi}(.) - \overline{\pi}^i_B(w_A) \end{aligned} \tag{5}$$

Note here that the status-quo profits $\overline{\pi}^{i\prime}_B(w_A)$ depends on w_A. Indeed, when R_i does not distribute good A, R_j still sells product A and faces the same w_A and therefore the demand for product B at retailer R_i still depends on w_A. The above program boils down to maximizing:

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

After reintegration of the downstream firms' FOCS given by (2) and using demand and equilibrium price symmetry, the first order condition of the above program is further simplified as follows when $w_{B1} = w_{B2} = w_A = c$:[10]

$$\begin{aligned} &2\frac{dp^*_{Ai}}{dw_A}\underbrace{[(p^* - c)(\frac{\partial D_{Aj}}{\partial p_{Ai}} + \frac{\partial D_{Bj}}{\partial p_{Ai}}) - (p^3_{Bj} - c)\frac{\partial D^3_{Bj}}{\partial p_{Ai}}]}_{\alpha} \\ &+2\frac{dp^*_{Bi}}{dw_A}\underbrace{[(p^* - c)(\frac{\partial D_{Aj}}{\partial p_{Bi}} + \frac{\partial D_{Bj}}{\partial p_{Bi}}) - (p^3_{Bj} - c)\frac{\partial D^3_{Bj}}{\partial p_{Bi}}]}_{\beta}. \end{aligned} \tag{6}$$

We have shown in lemma 1 that $\frac{dp^*_{Ai}}{dw_A} \geq 0$, and $|\frac{dp^*_{Bi}}{dw_A}| < |\frac{dp^*_{Ai}}{dw_A}|$. Note that the first term of β and α are identical due to demand symmetry; furthermore, as $|\frac{\partial D_{kj}}{\partial p_{li}}| \leq \frac{\partial D_{lj}}{\partial p_{li}}$, this term is positive. Furthermore, because of Assumption 2, the second term of β is more negative than the second term of α: hence $\beta \leq \alpha$. Therefore $\beta \geq 0$ is sufficient to have (6) > 0. Part (ii) of Assumption 3 ensures that the first order condition (6) is positive when $w_A = c$. This leads to the following lemma:

Lemma 6 *When discrimination is banned, under Assumption 3, the unique equilibrium unit wholesale price of product A is such that* $w^{nd}_A > c$.

Proof. See Appendix A.6. Part (ii) of Assumption 3 is sufficient to ensure that a unit price increase for product k at a retailer R_i leads to a larger increase in profit for the competitor R_j when he can sell the two products than when he sells only one product ($\frac{\partial \pi^i}{\partial p_{kj}} > \frac{\partial \pi^{i3}_k}{\partial p_{kj}}$). ■

All input prices being anticipated, we obtain the following lemma:

Lemma 7 *When discrimination is allowed, under Assumption 1-3, there exists a unique retail price equilibrium symmetric among retailers but asymmetric among products with* $\hat{p}_A = p^*_{ki}(w^{nd}_A, c, w^{nd}_A, c)$ *and* $\hat{p}_B = p^*_{ki}(c, w^{nd}_A, w^{nd}_A, c)$.

Proof. See Appendix A.4. ■

[10]See Appendix A.5 for more details.

3.5 The effect of removing the ban on input price discrimination on retail prices

Comparing lemmas 2 and 5 on the one hand, and lemmas 3 and 6 on the other hand, we obtain the following proposition:

Proposition 1 *Under Assumption 1-3, removing the ban on discrimination has no effect on the unit wholesale price of the private label products, $w_B^{nd} = c$; however it leads to a decrease in the unit wholesale price of the national brand product, $w_A^{nd} > c$.*

From Proposition 1 and lemma 1, we then obtain the following proposition:

Proposition 2 *Under Assumptions 1-4: Removing the ban on input price discrimination leads to a decrease in the retail price of product A: $p^* < \hat{p}_A$. Removing the ban may affect the price of product B in either direction, but this price variation is less pronounced, in absolute terms, than the price decrease for product A, $|p^* - \hat{p}_A| > |p^* - \hat{p}_B|$. When demand is linear, the price of product B remains unchanged ($\hat{p}_B = p^* < \hat{p}_A$).*

Proof. Proposition 2 results from lemma 1 and proposition 1. For the linear demand example, see Appendix A.7 ■

In what follows, we build on this theoretical analysis to define our empirical strategy. We use a difference-in-differences empirical analysis, building on Proposition 1 to define our treatment and comparison groups. We assess the effect of removing a ban on input price discrimination on retail prices, thus testing the predictions of Proposition 2.

4 A Natural Experiment: The French Grocery Sector

We briefly describe the main features of the French food retail sector and the change in the legislation on input price discrimination which provides us a natural experiment that enables to test our theoretical predictions.

4.1 The French Grocery Sector

The French grocery sector represented about 66% of total food purchases in 2015 with a retail network of 12 000 stores. It is highly concentrated with eight retail groups and a cumulative market share for the two first groups close to 42% in 2018 according to Kantar. Although each retail group gathers several retail chains, negotiations take place between the retail group, and sometimes alliances of retail groups, and their suppliers at a national level. On the suppliers' side, a few large groups, such as Danone Bonduelle or Lactalys, represent more than 40% of the total added value in the food chain, but 98% of suppliers in the agro-food industries are SMEs. Therefore, the balance of power between manufacturers and retailers is often in favor of the buyers and negotiations which take place annually from november to february are always a period of sharp tensions reported by the press.[11] These negotiations determine the tariff, i.e wholesale unit prices but also all types of unit rebates and fees such as slotting fees to obtain shelf space, or any fees in exchange 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 which is often used as a leverage in their negotiations. In France, the average market share of private label is about 32% in 2018[12]. Mostly, the retailer may buy the private label from small- and medium-sized dedicated firms or directly hold the production facilities. Second, the national brand producers themselves often supply the private- label goods to retailers. Finally, the retailer can also entrust the production of its private label to powerful manufacturers that, which have specialized in the production of private labels only and may work for several retailers at a time.[13]

4.2 The Legal Framework of Negotiations

In 1986, price discrimination by a supplier among similar buyers was forbidden in France. In particular, producers had to publish their general terms of sales that would be identical for all their "similar" buyers. In practice, this ban on input price discrimination ensures that two retailers obtained the same unit wholesale price from a supplier but they could still pay different fixed fees. Indeed, in practice, these fees gather a wide range of allowances such as slotting fees or commercial services, or deferred

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

[12]See Private Label Manufacturer Association.

[13]For instance, Richelieu Foods in the U.S., a private label food manufacturing company founded in 1862 that produces frozen pizza, salad dressings, sauces, marinades, condiments and deli salads to be marketed by other companies as their store brands, makes more than $200 million in yearly sales.

rebates, but their amount are quite difficult to assess. The opacity surrounding the negotiations would make it difficult for a court to establish that discriminatory fixed fees have been used.

Twenty years later, in 2008, the Loi de Modernisation Economique (LME) came back on this principle reauthorizing suppliers to price-discriminate between retailers. This reform, implemented on january, 1st 2009, 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 startedin 2005 to break-up the inflationary mechanism caused by the former legislation, the Galland Act. The 1996 Galland 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: both retailers and suppliers artificially increased the threshold thus binding final prices at the price-floor level and shared the profit through fixed fees.[14] In 2005, the Loi Dutreil started breaking-up this mechanism by enabling retailers to incorporate most of these rebates in the resale below-cost threshold. Specifically, all rebates representing more than 15% of the unit price invoiced should be accounted for to determine the new threshold. A last reform took place with Châtel Act in January 2008 enabling retailers to incorporate all types of rebates in the resale-below-cost law threshold. Although this reform took place almost simultaneously with the LME reform, we are confident that most of the effect of the Galland Act reform process came with the Dutreil Act of 2005 which was a quite radical reform.

5 Empirical Strategy and Sample Selection

Our objective is to estimate the price changes in food prices resulting from the authorization to price discriminate between manufacturers and retailers implemented in France in 2009.

[14]Biscourp et al. (2013) highlight that the correlation between local market concentration and retail prices collapsed after the Galland Act.

We first present in Section 5.1 our empirical strategy. Sections 5.2 presents our households data and explains how are constructed the variables of interest. Section 5.3 defines the affected and comparison groups while Section 5.4 details the criteria followed to construct our final dataset. Summary statistics are reported in Section 5.5. Finally, we provide graphical evidence of parallel price trends between the two groups for the pre-LME period in Section 5.6.

5.1 Empirical Strategy

A straightforward way to measure the price effect of the LME would consist in comparing the mean changes in prices for a group of products affected by the LME with the potential mean changes those products would have experienced under the ban. Because we cannot observe how prices of affected products would have evolved under the ban, we follow the program evaluation literature and compare the mean change in prices of affected products to the mean change in prices of "non affected" products, i.e. of a comparison group. In the theoretical approach presented in 3, we predicted that the LME must have had a differentiated effect regarding brand type (national brands *vs* private labels). Using these theoretical predictions, we are able to construct a valid counterfactual group and, and we implement a Difference-in-Differences (DID) approach.[15]

The principle of the double differencing is the following. The first difference allows to neutralize the difference in price level between products belonging to the affected and comparison groups. Such difference is likely to exist since private labels are on average less expensive than national brands. The second difference controls for temporal effects that could be otherwise confounded with the effect of the LME (cost or demand shocks concomitant with the LME). To illustrate, the LME was enacted at a time where a broad range of agricultural commodities (such as wheat, maize, rice and milk) experienced a sharp price increase (see European Commission, 2008). The passing of the law is then concomitant to an inflationary trend of food prices. Simply comparing the mean prices of affected products before and after the LME would generate estimates

[15]The DID estimation method was developed to conduct retrospective analyzes of policy outcomes. With the increasing availability of empirical data, it has been extensively used in IO to question various issues such as merger effects (e.g., Ashenfelter, Hosken, and Weinberg, 2015), vertical integration (e.g., Hastings and Gilbert, 2005), or switching costs (e.g., Viard, 2007).

biased upwards. However, by comparing the mean change in prices of affected products to the mean change in prices observed for the comparison group, we remove any unobserved time-variant factor that evenly affects both groups. It results the central identification assumption of our DID approach: we would obtain consistent estimates of the price effects of the LME under the assumption that, absent the law, the prices would have evolved identically between the affected and comparison groups.

5.2 Data and variable of interests

Household Scanner Data Our study uses household scanner data coming from the Kantar Worldpanel (2006-2010) survey. These data offer a precise and detailed representation of food shopping behaviors in France. The dataset records information on daily food purchases over a panel of more than 10,000 households residing in France, for a given year. The panel of households is representative of the French population. Purchase data are collected by the households themselves, usually by mean of a home scanner.[16] The households record information on the quantity and the expenditure for each product purchased, as well as the store type where the purchase was made (e.g., hypermarket, supermarket, specialized store) and, for retail chains, their name. Further, for products with a European Article Number (EAN, a 13-digit barcode wich is a superset of the Universal Product Code), the dataset contain detailed information on product characteristics. Hence a product can be described by up to 22 descriptive variables (such as flavor, container, and nutritional characteristics, for instance), plus the brand name and the name of the manufacturer.[17] Overall, each product can be clustered into more than 349 categories of food products which can themselves be aggregated into 61 families of products (see the classification in the Online Appendix). To illustrate, the product “Beurre Président Gastronomique, sweet butter, Normandy origin, aluminium packaging, 125g, 82% fat, without Omega 3 and cholesterol” is included in the product category “Butter” that belongs to the product family “Dairy products”.

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

[17] We remove fresh products such as fruits, vegetables, meat or fish because raw agricultural products were not concerned by the LME.

Variables of Interest Our identification strategy requires to determine the brand type of a product. However the Kantar Worldpanel data only inform about whether a product is a private label or not. We therefore adopt a more detailed classification and we classify each brand in 4 brand types following the usual classification in France: national brand (NB), private label (PL), discount private label (PL-D), and first-price brand (FP). The identification of first-price products (FP) is more complicated because some are sold under a specific store own-brand (e.g., Carrefour Discount) while the others are sold across several retailers under a generic name. To identify FP, we thus select products soldat a price below the annual average price of the product category minus twice the standard deviation. Remaining products are thus classified as national brands.[18]

The Kantar Worldpanel survey offers a unique opportunity to track product prices over time thanks to transaction data recorded by households. We define the price variable at the retail chain-product level since prices of all brand types are likely to be set at the chain level. Using the transaction data, we compute a mean unit price for each retail chain-product pair which summarizes the national price.[19] To ensure a sufficient number of purchase observations per retail chain-product pair, we aggregate the data over 4 weeks.[20] 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.[21]

5.3 The Affected and Comparison Groups

Following the predictions of our theoretical model, a natural definition of the affected and comparison groups can rest on the brand type of products. Because the lift of the ban on wholesale price discrimination only concerns products sold by a manufacturer to several retail groups, potentially all national brand products are affected by the LME.

[18] A detailed presentation of the classification method is provided in the Online Appendix.

[19] Data are trimmed to exclude transactions whose the standardized deviation of the unit price from the monthly average price is greater than 10 in absolute value.

[20] A 4-weeks period corresponds to a month in the terminology of Kantar. Hence, a year is composed of 13 periods of 4-weeks. Hereafter, we will use the term "month" to denote a 4-weeks period.

[21] Note that we use the purchase weights and the household weights provided by Kantar to correct the monthly purchased quantities and the monthly expenditures of a given product in such a way that they are representative of the French population.

However, to make sure that our affected group is well defined we keep only, among all NB products, those which are sold by competing retails groups.[22] In contrast, private labels, which often involve an exclusive partnership between a manufacturer and a retailer constitute natural candidates to the comparison group. Again, among PL products we select only those which indeed are only sold by one retail group. However, to obtain consistent estimates of the price effect of the LME, one must ensure that products belonging to the comparison group satisfy two requirements. First, as their prices should have evolved in the same way as those of the affected products absent the LME, it is important to select products that are exposed to similar demand and cost shocks and that are supposed to react in the same way (i.e. "common trends hypothesis"). Because PL compete with national brands within stores whereas PL-D compete with national brands between stores (conventional stores *vs* discount stores), we choose to restrict our attention to private labels only offered by conventional retailers (PL). With the same intention of comparing products exposed to similar demand shocks, we also remove from the comparison group all first-price products (FP).[23] Indeed, first-price products are low substitutes for national brands and then compete for different segments of the demand. Further, first-price products are characterized by higher pass-through than national brands, which is problematic in a context of a surge in commodities prices.

A second requirement is that products belonging to the comparison group are not affected by the lift of the ban. As shown by our theoretical model, this condition could not be guaranteed as it is only in the specification with linear demand that we predict PL prices are unchanged after authorizing input price discrimination. In the short run, the pricing of private labels may evolve in reaction to the change in prices of national brands and the model predicts that this variation can go either ways. This means that the estimate of the price effect of the LME could be biased upwards or downwards if private labels have been affected indirectly by the law, but in any case the sign of the effect could be reversed.

Given these considerations, our baseline definition of the affected group includes all national brand products sold in at least two retail groups, while the comparison

[22] As mentioned in Section 4.1, retailers bargain with supplier at the group level and not at the chain level.

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

group is composed of private labels offered by conventional retailers (i.e., PL).[24]

5.4 Sample Selection

Using data from the Kantar Worldpanel survey, we construct our sample based on 4 criteria that rely on: (i) the timing of the passing of the law, (ii) the distribution channels directly concerned by the law, (iii) the frequency of purchase observations, and (iv) the existence of an offer of private labels within a product category.

Time Period We consider the period that spans from 2006 to 2010, i.e. two and a half years before and two and a half years after the introduction of the LME. In order to eliminate the transitory period that has followed the introduction of the LME, we choose to remove the data corresponding to the six months following the introduction of the LME, that is from August to December 2008. In the robustness section, we test the sensitivity of our results to this time frame.

Distribution Channel We exclude from the sample food purchases made in distribution channels not directly targeted by the LME. Hence product purchases made in non-food distribution channels (e.g., gasoline stations, sporting goods retail chains), specialized distribution channels (e.g., farmer markets, frozen retail chains) as well as specialized shops (e.g., butchers, bakers, wine merchants) are removed from the dataset. After this selection, the data only contain food purchases made in food retail chains and their associated online food platforms (e.g. Chronodrive, Ooshop, or Télémarket), which corresponds to a total of 72 millions of transactions for the years 2006-2010.

Product We impose that each product must be purchased at least once per month over 36 months to be retained. Because a large proportion of products enters or exits during the five-year period of the study - only 5% of all products are purchased

[24]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 factors that could affect prices, such as product characteristics or purchase location, vary significantly across the affected and comparison groups, our estimate will be biased. For instance, if the affected group is only composed of national brands of organic dairy products whereas private labels that composed the comparison group correspond to low-quality and high-fat dairy products, one may concern that brand type is endogenous to price. This selection bias is however limited in our case because the LME has affected almost all product categories (exceptions are raw agricultural products negotiated on spot markets) and applied nationally.

each of the 59 months,[25] we have to account for the short shelf live of food products and we cannot impose that the final sample follows a balanced panel structure.[26] Our criterion selection results from a delicate tradeoff between integrating the largest number of products and ensuring that each product retained in the final sample is purchased both during the pre- and post-LME periods. Indeed, this last condition allows us to introduce product fixed-effects in the model regression and eliminate any sample composition effect that may bias the estimates. Overall, by defining a threshold of 36 months, we retain 84% of all products. A consequence of this selection criterion is however that seasonal products (such as Christmas turkey or Easter eggs, for instance) are excluded from the final sample.

Product Category Because all the product categories do not necessarily contain both national brands and private labels, we set a last criterion that requires that each product category retained must be present in both the affected and comparison groups. As before, the purpose of this selection criterion is to avoid any sample composition effect at the category level. Note that this criteria would de facto lead us to exclude of raw agricultural products for which there is no offer of private label products (such as baby food and drink or cheeses with designation of origin protected).[27]

5.5 Summary Statistics

Table 1 reports the composition of the final sample split by affected and comparison groups. The final sample is composed of 26,254 products split up in 168 product categories, which in turn are gathered into 57 product families. The allocation of products between the affected and comparison groups is relatively balanced, and on average a product category is composed of 74.21 products in the affected group and 82.06 products in the comparison group. Note that by construction the affected group contains only NB and the comparison group only PL. As expected, we observe that affected products are more expensive than comparison products. The average monthly unit

[25]Recall that we removed six months of data out of the 65 months due to the transitory period following the introduction of the LME.

[26]Another reason of the large proportion of product churning is the presence of low sales products like horsemeat for which we only observe few purchase observations per month. It is thus more likely for such product that no purchase is recorded for a given month.

[27]We provide in the Online Appendix the list of the product categories excluded after applying these two criteria.

Table 1: Summary Statistics for Affected and Comparison Products

	Affected group	Comparison group	Total
Panel A: Product			
Number of products	12,468	13,786	26,254
Number of product categories	168	168	168
Average number of products per category	74.21	82.06	156.27
Number of chain stores	53	41	55
Panel B: Brand type			
Percentage of national brand products	100	–	0.47
Percentage of private label products	–	100	0.53
Panel C: Price			
Mean of monthly average product price	10.27	7.11	9.33
S.D. of monthly average product price	49.67	19.24	43.01
Min. of monthly average product price	0.02	0.02	0.02
Max. of monthly average product price	4501.02	2364.53	4501.02
Panel D: Purchase transaction			
Number of purchase observations	14,904,852	10,785,417	25,690,269
Total expenditures	46,311,088	23,909,402	70,220,490

Notes: The table reports summary statistics on the composition of the affected and comparison groups as well as for the final sample. The statistics are calculated over the pre- and post-LME periods. The 6 months following the introduction of the LME are removed from the data.

price of national brand products is 10.27 while it barely exceeds 7 euros for private label products.

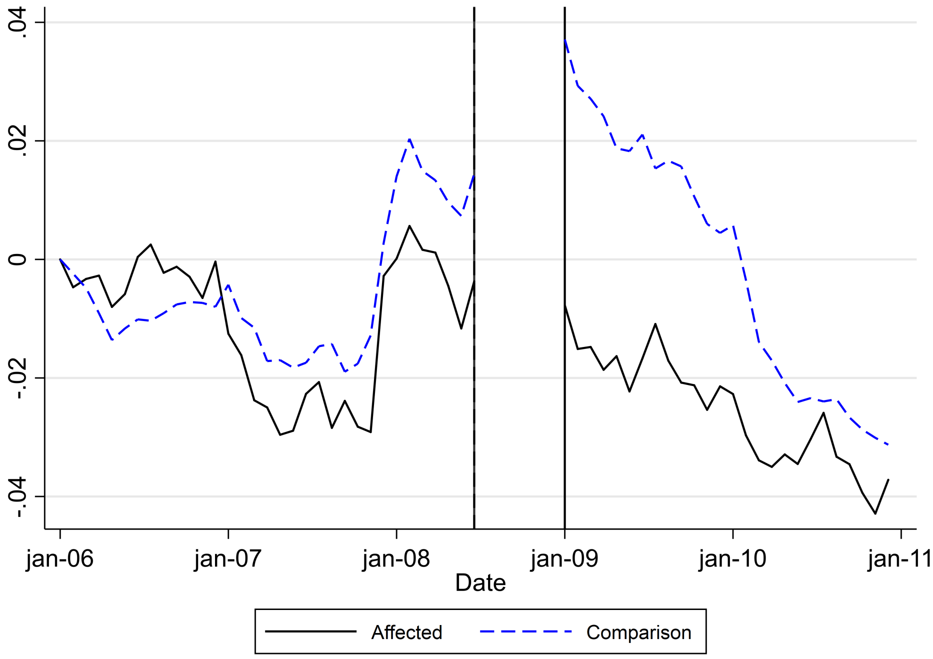
Overall, the monthly price data of products retained in the final sample are calculated based on 25,690,269 purchase observations. The most purchased product family is dairy products (13.44% of the observations), followed by cheese products (11.41%), cure products (7.24%), canned food (5.57%), bakery products (4.46%), and non-alcohol drinks (4.01%). In terms of expenditures, the final sample represents more than 70 million euros of cumulative food expenditures over the period 2006-2010, which is about 38% of the purchase expenditures recorded in the *Kantar Worldpanel* survey. This dataset then provides a unique opportunity to evaluate the price effect of intermediate price discrimination through a large-scale study.[28]

5.6 Price Trends

Our identification strategy relies on the assumption that, absent the LME, prices would have evolved identically between the affected and comparison groups. Although it is not possible to test directly this assumption, we follow standard practices in the litera-

[28] We provide additional summary statistics at the product category level in the Online Appendix.

Figure 2: Average Monthly Variations of Log Prices



Notes : The graph presents the average monthly variations of log prices for the affected (solid line) and comparison (dashed line) groups. Each point corresponds to an estimated coefficient $\widehat{\varphi}_l$. The first month is taken as reference. The time period between the vertical bars correspond to the 6 months following the introduction of the LME that are removed from the data.

ture and check whether the average prices of the affected and comparison groups follow similar trends in the pre-LME period (see, e.g., Angrist and Pischke, 2008; Ashenfelter, Hosken, and Weinberg, 2013).

We first provide graphical evidence of the "common trend" assumption by comparing the monthly average variations of prices for the affected and comparison groups. These monthly variations are obtained by regressing the weighted-average of (log) prices on a set of monthly dummy variables for each group 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. The 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 Fig.2 the monthly price variations $\widehat{\varphi}_l$ against the time period for the affected and comparison groups.

A first look at the figure shows that the average prices have evolved in a similar way for the affected and comparison groups during the pre-LME period. There is no obvious differential trend prior the passing of the LME, and both national brand and private label products seem to have experienced similar developments during this period. We also conduct a more formal analysis where we test the absence of a specific (linear) trend for the affected group prior to the LME. The null hypothesis cannot be rejected at the 10% level, which supports the "common trend" hypothesis (see Appendix C for further details).

Second, it is clear from Fig.2 that the average price of national brand products have decreased substantially after the introduction of the LME. This results in the short-term by an increase of the vertical gap between the lines of the affected and comparison groups. However, few months after the LME, prices of private label products have fallen even more than those of national brand products. At the end of 2010 both product types have experienced similar price decreases relative to the first month of 2006. In the next section, we test whether the differences in price changes between the affected and comparison groups are statistically significant.

6 Empirical Results

6.1 First Insights on Price Changes

We first compare how prices of all products included in the treatment and comparison groups have evolved before and after the introduction of the LME. The purpose is to have a basic idea of the price evolutions surrounding the passing of the law. To that, we estimate the following regression model with OLS using product level (log) prices as the dependent variable:

$$ln\left(P_{kit}\right) = \alpha + \beta PostLME_t + \mu_{ki} + \varepsilon_{kit} \tag{7}$$

where P_{kit} denotes the monthly average price for product-chain pair ki at month t, $PostLME_t$ is a dummy variable equal to one for months following the introduction of the LME, and μ_{ki} controls for product-chain specific fixed effects. The observations are weighted by the expenditure shares of food products, calculated at the national

Table 2: Prices Changes around the LME

Dependent variable: (log) of monthly average price		
Variable	(1)	(2)
$PostLME_t$	-0.0130***	
	(0.0029)	
$PostLME_t \times$ PL		0.0050***
		(0.0019)
$PostLME_t \times$ NB		-0.0144***
		(0.0032)
Chain-product FE	Yes	Yes
R^2	0.986	0.986
Observations	3050346	3050346

Notes : The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period.. The standard errors, shown in parentheses, are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

level during the pre-LME period. The estimates are reported in Table 2. Column (1) shows that prices decrease, on average, by 1.30% after the introduction of the LME. Column (2) splits the $PostLME_t$ variable by brand type in order to detect some specific trends. The estimate reveals that the price decrease observed after the LME is entirely driven by the change in price of national brand products. On the other hand, private label products have experienced, on average, a rise in price of 0.05%. These opposite price developments could be explained by different reactions to the introduction of the LME (as highlighted by our theoretical model) and/or to the exposure of external shocks. Assuming that both groups of products have been affected by identical shocks and react in a similar way, it is possible to infer a raw measure of the LME effect. By comparing the point estimates of each group of products, we find that authorizing intermediate price discrimination has decreased the price of national brand products by 1.94% compared to private label products.

6.2 Average Price Effect

Table 3 presents the (weighted-)estimates of the causal effect of the LME by comparing the mean change in prices that national brand products belonging to the affected group have occurred between the pre- and post-LME periods, to the mean change in prices of products included in the comparison group. To give more weight to products with higher sales, we weight the observations by the expenditure shares of food products, calculated at the national level during the pre-LME period. Column (1) first provides the estimation results of our baseline specification assuming that all products, both in the affected and the comparison groups, have experienced similar trends during the pre- and post-LME periods. This corresponds to estimate by OLS the following equation:

$$\ln(P_{kit}) = \alpha + \beta T_{ki} \times PostLME_t + \delta T_{ki} + \gamma PostLME_t + \mu_{ki} + \varepsilon_{kit} \qquad (8)$$

where P_{kit} denotes the monthly average price for product k in chain i at month t, T_{ki} is a dummy variable that characterizes the product-chain pair ki as belonging to the affected group, $PostLME_t$ is a dummy variable equal to one for months following the introduction of the LME, and μ_{ki} are product-chain fixed effects.

A concern with this baseline specification is that if a subgroup of products has experienced dissimilar trends or was exposed to unobserved transitory shocks at the time of the passing of the law, our estimates will be biased. In order to control for unobserved factors that could have differently affected our product sample, and that could be correlated with the effect of the law, we augment our baseline specification by two alternatives set of control variables/fixed-effects. We first consider the case of heterogenous trends among retail chains and introduce chain-month fixed effects. The idea is to control for chain-specific monthly shocks that could shift product prices evenly between the affected and the control groups and across stores of a chain. For instance, if a chain decided to cut down its prices drastically during the 2009 winter sales through important promotional activities, this could bias downward the estimated effect of the LME. In the same spirit, we consider the case of time-variant factors that could generate price changes for all products within a given category. We are particularly aware that the surge in agricultural commodities prices observed at the end of 2007 has differently impacted product prices regarding their category. We then introduce in a third specification category-month fixed effects to capture in a flexible way the category-specific monthly deviation of prices.

All these factors are supposed orthogonal to the LME and to affect products of the affected and comparison groups identically. Hence, they cannot explain some price differences between both groups. Consequently, the average price effect of the LME is captured by the coefficient β, and can be interpreted as the mean effect resulting from authorizing intermediate price discrimination on retail prices.

Table 3: Estimated Price Effect

Dependent variable: (log) price (P_{kit})			
		With monthly trend	
	Baseline	Chain	Category
Treatment × PostLME	-0.0195***	-0.0217***	-0.0336***
	(0.0037)	(0.0036)	(0.0031)
Chain-product FE	Yes	Yes	Yes
Chain-month FE	No	Yes	No
Category-month FE	No	No	Yes
R^2	0.986	0.986	0.987
Observations	3050346	3050173	3050338

Notes: The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The 6 months following the introduction of the LME are removed from the data. The standard errors, shown in parentheses, are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

Column (1) of Table 3 reports the baseline results of the estimation of Equation 8. Standard errors are clustered at the product level. Compared to private labels, we observe that the price of national brand products has significantly decreased after the introduction of the LME.[29] Adding monthly trends substantially strengthens the price decreasing effect of authorizing intermediate price discrimination (Columns 2-3). In particular, controlling for category-specific monthly deviations leads to a point estimate corresponding to a price decrease of -3.36% of national brand products relative to the price evolution of private label products (see Column 3). The point estimate difference with category monthly-trends or no trends highlights the importance to control for unobserved transitory shocks at the category level in our case.

The estimated causal effect reports a change of the *relative* price of national brand products (i.e., a change in prices of national brand products relative to price changes of products in the comparison group). To relate this finding to our theoretical result we further explore whether this relative variation in prices observed is indeed driven by a decrease in national brands prices and a less pronounced price variation (decrease or increase) in the price of private labels. Indeed, for instance in case both prices of national brands and private labels increase (but to a larger extent for the latter), we would observe similar relative price decrease in national brand with respect to private label due to the LME. A first step towards dismissing this scenario relies on the time-difference conducted in Column (2) of Table 2. The point estimates show that national brand products have experienced on average a significant price decrease after the LME,

[29]By running an unweighted regression model, we obtain a point estimate of -0.0114***(0.0015).

Table 4: Prices Changes around the LME

Dependent variable: $\lvert \widehat{P}_{ik}^{post} - \widehat{P}_{ik}^{pre} \rvert$		
	Unweighted	Weighted
Variable	(1)	(2)
T_{ik}	0.1067***	0.1555***
	(0.0176)	0.0132
Category FE	Yes	Yes
R^2	0.0681	0.0732
Observations	62028	62028

Notes : The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. Robust standard errors are showed in parentheses. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

whereas the prices of private labels significantly increased in the post-LME period. Next, we test more directly whether the change in prices, in absolute value, between the pre- and post-LME periods is higher for national brand products than for private labels products. To do so, we first compute the average price of a given product-chain pair *ki* over the pre- and post-LME periods. Then, we determine the change in price in absolute value and we regress this variable against the dummy variable T_{ki} as follows:

$$\lvert \widehat{P}_{ki}^{post} - \widehat{P}_{ki}^{pre} \rvert = T_{ki} + \eta_c + \varepsilon_{ki}$$

where η_c corresponds to product category fixed effects. The estimates, reported in Table 4, confirm that national brand products have experienced, on average, a larger change in price than that of private label products. This rules out two scenarii that would not be consistent with our theory but consistent with Table 3 result, namely (i) the national brand prices increase and the private label prices increase to a larger extent or (ii) the national brand prices decrease and the private label prices increase to a larger extent.

Our empirical analysis thus supports Proposition 2 that is, on average, authorizing intermediate price discrimination has caused a decrease in the prices of national brands and a variation in the prices of private labels of a smaller extent.

6.3 Heterogeneous Price Effect

Given the large number of dimensions covered by the final dataset, we further explore whether the price decreasing effect of authorizing input price discrimination differs among some of these dimensions. In what follows we analyse the heterogeneous effects of the law across retail chains and product categories.

Effects by retail chains We first report in Table 5 the estimation results when the average effect of the law is split by retailer. The regressions include category monthly trends. Columns (1)-(2) focus on the two main retailers of the French market, denoted R_1 and R_2, and show that they both have decreased their prices to a smaller extent than the other retailers. For instance, the market leader - R_1 - decreased its prices of national brands by 0.6 percentage point less than its rivals. In Column (3), we decompose the effect for the seven largest retail groups, denote R_1 to R_7, and we observe significant differences across retailers. While all retailers have dropped the relative price of national brand products by at least 2%, we find that retailers R_5 to R_7 have experienced a larger decrease slightly higher than 4%.

Two remarks are in order. First, it appears that the implementation of the LME has significantly boosted intra-brand competition. The law has achieved its goal to restore an effective price competition in the French food retail sector after a long inflationary period initiated by the adoption of the Galland Act (see Biscourp, Boutin, and Vergé, 2013). Second, it is interesting to note that the three retailers that have experienced the largest drop in prices are among the smallest in terms of market share and also among the most expensive. This result suggests that the lift of the ban has forced high-price retailers to cut down drastically prices of branded products in order not to weaken their market share position. This differentiated price effect has then contributed to price convergence of national brand products after the implementation of the law.

Effect by product category Substitutions between national brands and private labels vary a lot from one category of product to another. Therefore, it is likely that the price effects of authorizing a ban on input prices will affect difefrently the relative decrease in national brands retail prices. In order to explore this dimension, we estimate Equation 8 for each product category separately, and we replace the category-month fixed effects

Table 5: Estimated Price Effect by Retailer

Dependent variable: (log) price (P_{kit})			
	(1)	(2)	(3)
Treatment × PostLME	-0.0344***	-0.0352***	-0.0198**
	(0.0031)	(0.0031)	(0.0090)
Treatment × PostLME × R_1	0.0061***		-0.0083
	(0.0015)		(0.0085)
Treatment × PostLME × R_2		0.0076***	-0.0076
		(0.0023)	(0.0087)
Treatment × PostLME × R_3			-0.0110
			(0.0087)
Treatment × PostLME × R_4			-0.0122
			(0.0080)
Treatment × PostLME × R_5			-0.0203**
			(0.0091)
Treatment × PostLME × R_6			-0.0219**
			(0.0087)
Treatment × PostLME × R_7			-0.0214**
			(0.0088)
Chain-product FE	Yes	Yes	Yes
Category-month FE	Yes	Yes	Yes
R^2	0.987	0.987	0.987
Observations	3050338	3050338	3050338

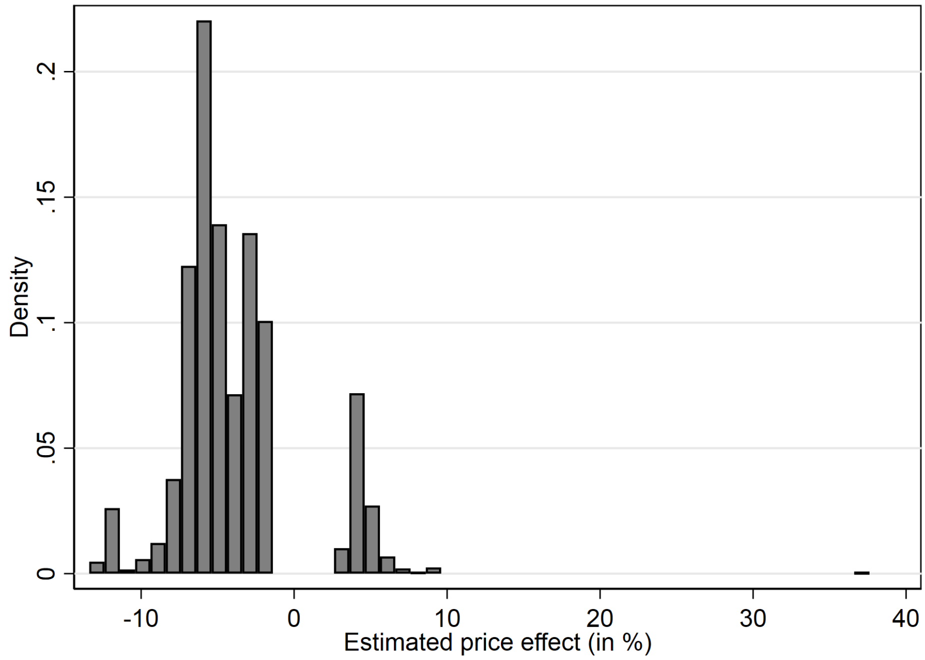
Notes: The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The 6 months following the introduction of the LME are removed from the data. The standard errors, shown in parentheses, are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

by monthly dummies. We present graphically the distribution of the estimated price effects in Fig.3. The estimated price effects not statistically significant are excluded from the analysis.

We observe that a large proportion of product categories have experienced a price decrease – 82% of the product categories – which reinforces the robustness of the previous findings. The estimated price effects range from -13.39% to +37.06%. The three largest decreases are mustard (-13.39%), semolina & polenta (-13.09%), and fresh sausages (-11.54%), whereas frozen fruit juices (+37.06%), Cider (+9.56%) and Mimolette (cheese) (+8.14%) have known significant price increase. All these product categories are however marginal in terms of sales, and the bulk of products have experienced a fall in prices between 0 to 1% due to the passing of the LME.

We then investigate whether variations in the estimated price effects can be explained by the bargaining power of suppliers. Since the bargaining power of suppliers cannot be observed directly, we use as a proxy the market share of private label products within a product category. The rationale being that, when input price discrimi-

Figure 3: Distribution of Estimated Price Effects



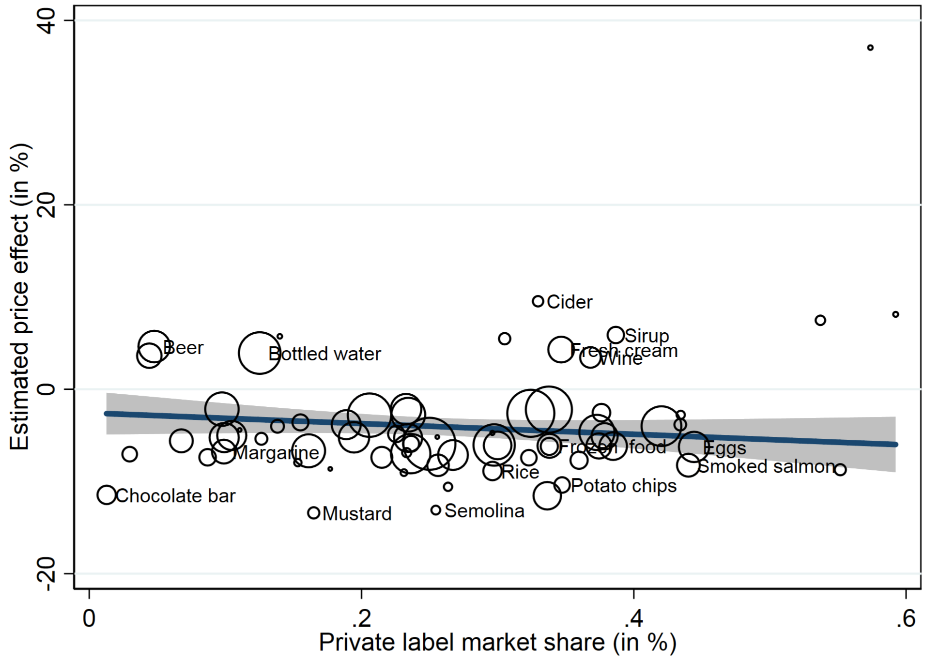
Notes: The graph presents the distribution of the price effects, estimated at the product category level, that are statistically significant. The observations are weighted by the expenditure shares of product categories, calculated at the national level during the pre-LME period.

nation is banned, national brand suppliers facing low competition from private labels are expected to impose higher wholesale prices. Hence, the price decrease of national brands resulting from the authorization of input price discrimination is expected to be lower for product categories in which private labels have a low market share. We plot in Fig.4 the price effect estimated for each product category against the market share of private labels within the product category. It appears from the scatter plot that no significant correlation exists between the estimated price effect and the market share of private labels; and this is confirmed by running a OLS regression. It results that the change in prices of national brands were not driven by the competition intensity exerted by private labels.

7 Robustness Checks

We conduct robustness checks in order to test the sensitivity of our result with respect to two hypotheses of the analysis. First, we examine in Section 7.1 how the result varies when adopting alternative definitions of the comparison group. Then, we focus in Section 7.2 on the definition of the time period of the study.

Figure 4: Private Labels Market Shares and Estimated Price Effects



Notes: The graph plots the average price effect estimated at the product category level against the market shares of private labels calculated by product category. The solid line corresponds to the linear regression line and 10% confidence bands are represented by the shaded area. The symbol weights are proportional to the expenditure shares of product categories over the total food expenditures.

7.1 Alternative Definition of the Comparison Group

We consider two alternative definitions of the comparison group. A first variant of the comparison group consists in considering exclusively private labels products offered by discounters (i.e., PL-D). Similarly to conventional retailers, discounters' private labels are supposed unaffected by the authorization to wholesale price discriminate. Moreover, prices of discounters' private labels are more likely to have been less impacted by a change in prices of national brands than private labels offered by conventional retailers due to a lower degree of substitutability. A second variant of the comparison group encompasses previous definitions and accounts for all private labels (i.e., PL and PL-D) as well as all first-price products sold under retailers own-brand (i.e., FP-PL).

We report in Table 6 the estimates of Equation 8 using both alternative definitions of the comparison group as well as the estimate with the baseline definition in the first row for ease of comparison. Looking at the estimated price effect when using discounters' private labels (PL-D) as the comparison group, we still find a price decreasing effect of authorizing input price discrimination for national brands. As expected, the

Table 6: Alternative Definitions of the Comparison Group

	$\widehat{\beta}$			
Comparison group	Coef.	S. E.	Obs.	R^2
Baseline (PL)	-0.0336***	0.0031	3,050,338	0.9870
PL-D	-0.0437***	0.0041	2,259,826	0.9940
PL & FP-PL	-0.0393***	0.0034	3,204,530	0.9870
PL, FP-PL & PL-D	-0.0401***	0.0031	3,483,154	0.9873

Notes: The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The 6 months following the introduction of the LME are removed from the data. The standard errors (denoted S. E.) are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

estimated price effect is lower in this case, compared to the one obtained with the baseline definition, because changes in prices of discounters' private labels are supposed more disconnected to the fall in prices of national brands. Finally, when the comparison group includes all private labels and all first-price products sold under retailers own-brand (i.e., PL, FP-PL & PL-D), we estimate a price effect close to the one obtained with the baseline definition.

7.2 Transitory Period

One difficulty of the retrospective analysis of the effect of the LME is to determine from which period suppliers have started to price discriminate across retailers. In the baseline analysis, we assume that suppliers have begun to price discriminate from the first annual price negotiations following the introduction of the LME. These annual price negotiations take usually place from November to the end of February, but some retailers have decided in the past to close the negotiations at the end of the civil year. As a result, we have decided to remove the period between August and December (i.e., 6 months) in order to exclude this transitory period. However, it is possible that the application of the LME has taken more time. We then examine whether accounting for a longer transitory period changes our result. As a robustness check, we define a transitory period that begins in August 2008 and ends 18 months later. We report in the second row of Table 7 the estimate obtained. As before, we report in the first row the baseline estimate for ease of comparison. We find that our result holds when defining a longer transitory period. Though the price effect is almost twice lower we still find that the relative prices of national brands have substantially decreased after

Table 7: Alternative Time Frames

Transitory period	$\widehat{\beta}$ Coef.	S. E.	Obs.	R^2
Baseline (2008/08–2008/12)	-0.0336***	0.0031	3,050,338	0.9870
2008/01–2008/12	-0.0368***	0.0033	2,650,216	0.9866
2008/08–2009/06	-0.0288***	0.0032	2,656,826	0.9869
2006/01–2006/12 & 2008/08–2008/12	-0.0243***	0.0029	2,443,356	0.9873

Notes: The observations are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The comparison group is defined as in the baseline specification. The standard errors (denoted S. E.) are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

authorizing input price discrimination.[30]

Finally, in a last robustness test, we study whether our result is sensitive when including the first year of the period of the study, because we have seen that the price trends were not perfectly parallel between the affected and comparison groups during this year. We then exclude the year 2006 from the analysis and we re-estimate Equation 8. The estimate of the price effect is reported in the last row of Table 7, and it appears that our finding is slightly affected when choosing a different time frame.

7.3 Price effect through time

We also have run regressions, keeping the same transitory period 01/08/2008 to 01/12/2008 as in our initial analysis, but spliting the effect in the POST LME period between the year 2009 (year 1) and the year 2010 (year 2). Table 8 shows that there is still a significant and negative impact of the LME on national brand prices but lower in the second than in the first year after LME.

8 Conclusion

This paper takes advantage of a reform, named Loi de Modernisation de l'Economie, that took place in 2008 and repealed the non-discrimination on input prices principle which has been in force since 1986 into the French retail industry. This natural experiment provides a unique opportunity to assess ex-post the change in retail prices resulting from authorizing input price discrimination.

We carry out the ex-post evaluation thanks to the availability of household scanner

[30]Part of the difference in the estimated price effects is explained by changes in the product sample.

Table 8: Estimated Price Effect

Dependent variable: (log) price (P_{kit})			
		With monthly trend	
	Baseline	Chain	Category
Treatment × Year 1	-0.0120***	-0.0333***	-0.0438***
	(0.0036)	(0.0034)	(0.0032)
Treatment × Year 2	-0.0277***	-0.0095**	-0.0227***
	(0.0039)	(0.0040)	(0.0032)
Chain-product FE	Yes	Yes	Yes
Chain-month FE	No	Yes	No
Category-month FE	No	No	Yes
R^2	0.986	0.986	0.987
Observations	3050346	3050173	3050338

Notes: The observations in Column (2) are weighted by the expenditure shares of food products, calculated at the national level during the pre-LME period. The 6 months following the introduction of the LME are removed from the data. The standard errors, shown in parentheses, are clustered at the product level. *, **, *** indicate significance at the 10%, 5% et 1% level, respectively.

data that contain detailed information on food purchases in supermarkets over the period 2006-2010. One appeal of the data is its scope which allows us to provide a large-scale evaluation of the reform on the basis of thousands of food products. We ground our empirical strategy on the predictions derived from an original model which enables us to highlight the economic forces at play. The model depicts vertically related markets in which, at the upstream level, a supplier offers its good to two multi-product retailers (national brand supplier) whereas two other producers (private label suppliers) have each an exclusive dealing relationship with one retailer. Exclusive dealing producers cannot price discriminate by assumption and therefore are not directly affected by the law. In contrast the supplier who sells to competing retailers is directly affected by the law. We show that, with non linear secret contracts, removing the ban on input price discrimination decreases the price of the national brand and has an indirect lower extent effect on private label retail prices.

Based on these predictions, we conduct a difference-in-differences analysis by comparing the mean change in prices of national brands to that of private labels. Our results show that removing the ban on input price discrimination indeed led to a decrease in national brands retail prices by 3.36% after the reform, relative to products belonging to the comparison group. Further, we show that the price decreasing effect of authorizing input price discrimination is observed in almost all product categories (82%).

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Appendix

A The Model: Appendix

A.1 Well-behaved reaction functions

We determine here the condtions under which the retailer's reaction functions are well-behaved. As retailers and products are assumed to be substitutes, we want to ensure that $\frac{dp^r_{Ai}}{dp_{Aj}} \geq 0$ and $\frac{dp^r_{Bi}}{dp_{Bj}} \geq 0$: if a retailer increases the price of product $K \in A, B$, everything else being equal, then its competitor shall react by increasing the price of product $K \in A, B$ also.

Let $(p_{Ai}, p_{Bi}, p_{Aj}, p_{Bj})$ denote the vector of prices. The first order conditions for retailer $i = 1, 2$ are written as follows:

$$\begin{aligned} \pi^i_1 &= 0 \\ \pi^i_2 &= 0 \end{aligned} \tag{9}$$

Solving these two first order conditions, for a given w_{Ai} and w_{Bi} determine the best response final prices p^r_{Ai} and p^r_{Bi} as a function of the final prices of the rival retailer.

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, we have

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

We have assumed $\Delta \geq 0$ to ensure concavity of the retailers profit functions. In order to ensure that the prices of each products at the two retailers are strategic complements, we make the following assumption:

$$\pi^i_{12}\pi^i_{23} - \pi^i_{22}\pi^i_{13} \geq 0$$
$$\pi^i_{21}\pi^i_{14} - \pi^i_{11}\pi^i_{24} \geq 0$$

Finally, we consider the variation of the best response prices with respect to their unit costs w_{Ai} and w_{Bi}. Consider retailer R_i. For a given p_{Aj}, p_{Bj} we have:

$$\frac{dp^r_{Ai}(p_{Aj}, p_{Bj})}{dw_{Ai}} = \frac{\pi^i_{22}\frac{\partial D_{Ai}}{\partial p_{Ai}} - \pi^i_{12}\frac{\partial D_{Ai}}{\partial p_{Bi}}}{\Delta}$$
$$\frac{dp^r_{Bi}(p_{Aj}, p_{Bj})}{dw_{Ai}} = \frac{\pi^i_{11}\frac{\partial D_{Ai}}{\partial p_{Bi}} - \pi^i_{21}\frac{\partial D_{Ai}}{\partial p_{Ai}}}{\Delta}$$

and similarly

$$\frac{dp^r_{Ai}(p_{Aj}, p_{Bj})}{dw_{Bi}} = \frac{\pi^i_{22}\frac{\partial D_{Bi}}{\partial p_{Ai}} - \pi^i_{12}\frac{\partial D_{Bi}}{\partial p_{Bi}}}{\Delta}$$
$$\frac{dp^r_{Bi}(p_{Aj}, p_{Bj})}{dw_{Bi}} = \frac{\pi^i_{11}\frac{\partial D_{Bi}}{\partial p_{Bi}} - \pi^i_{21}\frac{\partial D_{Bi}}{\partial p_{Ai}}}{\Delta}$$

A sufficient condition to ensure that the best response price for each good increases with the cost of this good (that is, $\frac{dp^r_{Ai}(p_{Aj},p_{Bj})}{dw_{Ai}} \geq 0$ and $\frac{dp^r_{Bi}(p_{Aj},p_{Bj})}{dw_{Bi}} \geq 0$) is:

$$|\pi^i_{11}| \geq |\pi^i_{21}|$$
$$|\pi^i_{22}| \geq |\pi^i_{12}|$$

Note that assumption 2 implies $\Delta \geq 0$.

A.2 Proof of lemma1

Let $p_{ki} = (p_{ki}, p_{li}, p_{kj}, p_{lj})$ denote the vector of prices. We denote retailer i's profit by $\pi^i = (p_{Ai} - w_A)D_{Ai}(.) + (p_{Bi} - c)D_{Bi}(.)$. We note f_k the derivative of function f wrt.

the k_{th} argument.

The first order conditions (10) can thus be written for $i = 1, 2$ as:

$$\pi_1^i = 0 \quad (10)$$
$$\pi_2^i = 0 \quad (11)$$

By totally differentiating the first order condition (10), for $i = 1$ say, with respect to w_A, we obtain:

$$\frac{dp_{A1}^*}{dw_A} = -\frac{-\frac{\partial D_{A1}}{\partial p_{A1}} + \pi_{12}^i \frac{dp_{B1}}{dw_A} + \pi_{13}^i \frac{dp_{A2}}{dw_A} + \pi_{14}^i \frac{dp_{B2}}{dw_A}}{\pi_{11}^i}$$
$$\frac{dp_{B1}^*}{dw_A} = -\frac{-\frac{\partial D_{A1}}{\partial p_{B1}} + \pi_{21}^i \frac{dp_{A1}}{dw_A} + \pi_{23}^i \frac{dp_{A2}}{dw_A} + \pi_{24}^i \frac{dp_{B2}}{dw_A}}{\pi_{22}^i}$$

Because firms 1 and 2 are perfectly symmetric, we have $\frac{dp_{A1}}{dw_A} = \frac{dp_{A2}}{dw_A}$ and $\frac{dp_{B1}}{dw_A} = \frac{dp_{B2}}{dw_A}$. Solving this system of equations, we obtain :

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

where

$$W = \frac{\partial D_{A1}}{\partial p_{B1}} \left[\pi_{12}^i + \pi_{14}^i\right] > 0$$
$$X = \frac{\partial D_{A1}}{\partial p_{A1}} \left[\pi_{24}^i + \pi_{22}^i\right] > 0$$
$$Y = \frac{\partial D_{A1}}{\partial p_{A1}} \left[\pi_{21}^i + \pi_{23}^i\right] < 0$$
$$Z = \frac{\partial D_{A1}}{\partial p_{B1}} \left[\pi_{13}^i + \pi_{11}^i\right] < 0$$
$$D = (\pi_{12}^i + \pi_{14}^i)(\pi_{21}^i + \pi_{23}^i)$$
$$E = (\pi_{13}^i + \pi_{11}^i)(\pi_{24}^i + \pi_{22}^i).$$

The signs of W, X, Y and Z come from Assumption 3. Assume that the following assumption 4 is satisfied:

$$\begin{aligned}
|\pi^i_{13} + \pi^i_{11}| &> \pi^i_{23} + \pi^i_{21} \\
|\pi^i_{24} + \pi^i_{22}| &> \pi^i_{14} + \pi^i_{12}
\end{aligned}$$

This assumption 4 implies that an infinitesimal increase in p_{ki} and p_{kj} affects more the marginal profit of a retailer on product k than the marginal profit of a retailer on product l. First, this assumption is sufficient to ensure that $E - D > 0$, $X - W > 0$ and thus that $\frac{dp^*_{Ai}}{dw_A} > 0$.

We now prove that it is also sufficient to ensure that: $|\frac{dp^*_{Ai}}{dw_A}| > |\frac{dp^*_{Bi}}{dw_A}|$.

- Assume first that $Z - Y < 0$ then we $|\frac{dp^*_{Ai}}{dw_A}| > |\frac{dp^*_{Bi}}{dw_A}|$ if and only if $X - W > Y - Z$ which rewrites as:

$$\frac{\partial D_{A1}}{\partial p_{A1}}(\pi^i_{24} + \pi^i_{22} - \pi^i_{23} - \pi^i_{21}) > \frac{\partial D_{A1}}{\partial p_{B1}}(\pi^i_{12} + \pi^i_{14} - \pi^i_{13} - \pi^i_{11})$$

which, given the sign of each term can be rewritten as:

$$|\frac{\partial D_{A1}}{\partial p_{A1}}|(\pi^i_{23} + \pi^i_{21} - \pi^i_{24} - \pi^i_{22}) > \frac{\partial D_{A1}}{\partial p_{B1}}(\pi^i_{12} + \pi^i_{14} - \pi^i_{13} - \pi^i_{11}) \tag{12}$$

$$|\frac{\partial D_{A1}}{\partial p_{A1}}|(|\pi^i_{24} + \pi^i_{22}| + \pi^i_{23} + \pi^i_{21}) > \frac{\partial D_{A1}}{\partial p_{B1}}(|\pi^i_{13} + \pi^i_{11}| + \pi^i_{12} + \pi^i_{14}) \tag{13}$$

and therefore, the inequality 13 is always satisfied at the neighbourhood of the symmetric price equilibrium because $|\frac{\partial D_{A1}}{\partial p_{A1}}| > \frac{\partial D_{A1}}{\partial p_{B1}}$ and $\pi^i_{12} = \pi^i_{21}$,$\pi^i_{23} = \pi^i_{14}$,$\pi^i_{11} = \pi^i_{12}$, $\pi^i_{13} = \pi^i_{24}$.

- Assume that $W - Y > 0$, then we have $|\frac{dp^*_{Ai}}{dw_A}| > |\frac{dp^*_{Bi}}{dw_A}|$ if and only if $X - W > Z - Y$ which rewrites as:

$$\frac{\partial D_{A1}}{\partial p_{A1}}(\pi^i_{24} + \pi^i_{22} + \pi^i_{23} + \pi^i_{21}) > \frac{\partial D_{A1}}{\partial p_{B1}}(\pi^i_{12} + \pi^i_{14} + \pi^i_{13} + \pi^i_{11})$$

which can be rewritten as :

$$\frac{\partial D_{A1}}{\partial p_{A1}}(\pi^i_{23} + \pi^i_{21} - |\pi^i_{24} + \pi^i_{22}|) > \frac{\partial D_{A1}}{\partial p_{B1}}(\pi^i_{12} + \pi^i_{14} - |\pi^i_{13} + \pi^i_{11}|) \tag{14}$$

The inequality (14) always holds under assumption 4 and because $|\frac{\partial D_{A1}}{\partial p_{A1}}| > \frac{\partial D_{A1}}{\partial p_{B1}}$

and at the neighbourhood of the symmetric price equilibrium (where $\pi^i_{12} = \pi^i_{21}, \pi^i_{23} = \pi^i_{14}, \pi^i_{11} = \pi^i_{12}, \pi^i_{13} = \pi^i_{24}$) .

A.3 Proof of lemma 2

Using symmetry, the FOCs simplified as follows:

$$0 = (w_{Bi} - c)\frac{\partial D_{Bi}}{\partial p_{Bi}}\frac{dp^r_{Bi}}{dw_{Bi}} + (w_{Bi} - c)\frac{\partial D_{Bi}}{\partial p_{Ai}}\frac{dp^r_{Ai}}{dw_{Bi}}.$$

Computing the best-response pass-through, we obtain:

$$\frac{dp^r_{Ai}}{dw_{Bi}} = -\frac{-\frac{\partial D_{Bi}}{\partial p_{Ai}} + \pi^i_{12}\frac{dp^r_{Bi}}{dw_{Bi}}}{\pi^i_{11}}, \quad \frac{dp^r_{Bi}}{dw_{Bi}} = -\frac{-\frac{\partial D_{Bi}}{\partial p_{Bi}} + \pi^i_{21}\frac{dp^r_{Ai}}{dw_{Bi}}}{\pi^i_{22}}$$

Therefore:

$$\frac{dp^r_{Ai}}{dw_{Bi}} = \frac{-\frac{\partial D_{Bi}}{\partial p_{Bi}}\pi^i_{12} + \frac{\partial D_{Bi}}{\partial p_{Ai}}\pi^i_{22}}{\Delta}, \quad \frac{dp^r_{Bi}}{dw_{Bi}} = \frac{\pi^i_{11}\frac{\partial D_{Bi}}{\partial p_{Bi}} - \pi^i_{21}\frac{\partial D_{Bi}}{\partial p_{Ai}}}{\Delta}$$

where $\Delta = \pi^i_{11}\pi^i_{22} - \pi^i_{12}\pi^i_{21} > 0$. Because of Assumption 1 and 2, we have $\frac{dp^r_{Bi}}{dw_{Bi}} > 0$ and $\frac{dp^r_{Bi}}{dw_{Bi}} > |\frac{dp^r_{Ai}}{dw_{Bi}}|$ for $i = 1, 2$. It results that $|\frac{\partial D_{Bi}}{\partial p_{Bi}}\frac{dp^r_{Bi}}{dw_{Bi}}| > |\frac{\partial D_{Bi}}{\partial p_{Ai}}\frac{dp^r_{Ai}}{dw_{Bi}}|$, and therefore $w_{Bi} = c$ is the unique symmetric equilibrium.

A.4 Proof of lemma 3

U_A maximizes the following profit:

$$\begin{aligned}
\underset{w_{Ai}, F_{Ai}}{Max}\, & (w_{Ai} - c)D_{Ai}(p^r_{Ai}, p^r_{Bi}, p^*_{Aj}(w^d), p^*_{Bj}(w^d)) + F_{Ai} \\
& + (w^d_{Aj} - c)D_{Aj}(p^*_{Aj}(w^d), p^*_{Bj}(w^d), p^r_{Ai}, p^r_{Bi}) + F^d_{Aj} \\
s.t. F_{Ai} &= (p^r_{Ai} - w_{Ai})D_{Ai}(.) + (p^r_{Bi} - w^d_{Bi})D_{Bi}(.) - \overline{\pi}^i_B
\end{aligned} \tag{15}$$

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 the supplier offers retailer R_i the fixed fee that satisfies the retailer's participation constraint, that is, $F_{Ai} = (p^r_{Ai} - w_{Ai})D_{Ai}(.) + (p^r_{Bi} - w^d_{Bi})D_{Bi}(.) - \overline{\pi}^i_B$, and sets each unit price w_{Ai} to maximize:

$$\underset{w_{Ai}}{Max}(p^r_{Ai}-c)D_{Ai}(.)+(p^r_{Bi}-w^d_{Bi})D_{Bi}(.)+(w^d_{Aj}-c)D_{Aj}(.)+F^d_{Aj} \tag{16}$$

$$0=\tfrac{dp^r_{A1}}{dw_{A1}}[(p^r_{A1}-c)\tfrac{\partial D_{A1}}{\partial p_{A1}}+(p^r_{B1}-w^d_{B1})\tfrac{\partial D_{B1}}{\partial p_{A1}}+(w^d_{A2}-c)\tfrac{\partial D_{A2}}{\partial p_{A1}}+D_{A1}]+$$
$$\tfrac{dp^r_{B1}}{dw_{A1}}[(p^r_{A1}-c)\tfrac{\partial D_{A1}}{\partial p_{B1}}+(p^r_{B1}-w^d_{B1})\tfrac{\partial D_{B1}}{\partial p_{B1}}+(w^d_{A2}-c)\tfrac{\partial D_{A2}}{\partial p_{B1}}+D_{B1}].$$

Simplifying yields the following condition:

$$0=\tfrac{dp^r_{A1}}{dw_{A1}}[(w_{A1}-c)\tfrac{\partial D_{A1}}{\partial p_{A1}}+(w^d_{A2}-c)\tfrac{\partial D_{A2}}{\partial p_{A1}}]+$$
$$\tfrac{dp^r_{B1}}{dw_{A1}}[(w_{A1}-c)\tfrac{\partial D_{A1}}{\partial p_{B1}}+(w^d_{A2}-c)\tfrac{\partial D_{A2}}{\partial p_{B1}}].$$

Obviously there exists a symmetric equilibrium where $w_{A1}=w_{A2}=c$.

Is this equilibrium unique? Focusing on symmetric equilibria, we can rewrite this condition as:

$$0=(w_{Ai}-c)[\tfrac{dp^r_{A1}}{dw_{A1}}\underbrace{(\tfrac{\partial D_{A1}}{\partial p_{A1}}+\tfrac{\partial D_{A2}}{\partial p_{A1}})}_{(i)}+\tfrac{dp^r_{B1}}{dw_{A1}}\underbrace{(\tfrac{\partial D_{A1}}{\partial p_{B1}}+\tfrac{\partial D_{A2}}{\partial p_{B1}})}_{(ii)}].$$

The term (i) is negative under Assumption (1) and (ii) is positive under Assumption (1). If $\frac{dp^r_{B1}}{dw_{A1}}<0$, it is immediate that the unique symmetric equilibrium is such that $w_{Ai}=c$. If $\frac{dp^r_{B1}}{dw_{A1}}>0$, under Assumption 2, the symmetric equilibrium in which $w_{Ai}=c$ is the only symmetric equilibrium, because $|\frac{dp^r_{A1}}{dw_{A1}}|>|\frac{dp^r_{B1}}{dw_{A1}}|$ and $|(i)|>|(ii)|$.

A.5 Proof of Proposition

The above programme boils down to maximizing:

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

After simplification and reintegration of the downstream firms' FOCS given by (2), and using $w_{B1}^{nd} = w_{B2}^{nd} = c$, the first order condition of this program becomes:

$$\begin{aligned}0 = &\frac{dp_{A1}^*}{dw_A}[(w_A - c)\frac{\partial D_{A1}}{\partial p_{A1}} + (p_{A2}^* - c)\frac{\partial D_{A2}}{\partial p_{A1}} + (p_{B2}^* - c)\frac{\partial D_{B2}}{\partial p_{A1}}]\\ &+\frac{dp_{B1}^*}{dw_A}[(w_A - c)\frac{\partial D_{A1}}{\partial p_{B1}} + (p_{A2}^* - c)\frac{\partial D_{A2}}{\partial p_{B1}} + (p_{B2}^* - c)\frac{\partial D_{B2}}{\partial p_{B1}}]\\ &+\frac{dp_{A2}^*}{dw_A}[(w_A - c)\frac{\partial D_{A2}}{\partial p_{A2}} + (p_{A1}^* - c)\frac{\partial D_{A1}}{\partial p_{A2}} + (p_{B1}^* - c)\frac{\partial D_{B1}}{\partial p_{A2}}]\\ &+\frac{dp_{B2}^*}{dw_A}[(w_A - c)\frac{\partial D_{A2}}{\partial p_{B2}} + (p_{A1}^* - c)\frac{\partial D_{A1}}{\partial p_{B2}} + (p_{B1}^* - c)\frac{\partial D_{B1}}{\partial p_{B2}}]\\ &-\frac{\partial \overline{\pi}_B^1(w_A)}{\partial w_A} - \frac{\partial \overline{\pi}_B^2(w_A)}{\partial w_A}.\end{aligned}$$

$$\overline{\pi}_B^i(w_A) = (p_{Bi}^3 - c)D_{Bi}^3(p_{Bi}^3, p_{Aj}^*, p_{Bj}^*) \tag{17}$$

As we assume that there is no interim observability of wholesale prices and of acceptance or rejection of offers, only the price p_{Bi}^3 is adjusted because R_i is able to adapt its price but not the rival R_j who is not aware that R_i rejected the offer. Morever, we know that p_{Bi}^3 is such that:

$$(p_{Bi}^3 - c)\frac{\partial D_{Bi}^3}{\partial p_{Bi}^3} + D_{Bi}^3 = 0 \tag{18}$$

Therefore we have:

$$\frac{\partial \overline{\pi}_B^i(w_A)}{\partial w_A} = (p_{Bi}^3 - c)(\frac{\partial D_{Bi}^3}{\partial p_{Aj}}\frac{p_{Aj}^*}{\partial w_A} + \frac{\partial D_{Bi}^3}{\partial p_{Bj}}\frac{p_{Bj}^*}{\partial w_A}) \tag{19}$$

Simplifying the expression for $w_A = c$ and using the corresponding retail price equilibrium $p_{Ai}^* = p_{Bi}^* = p^*$, the expression further simplifies as:

$$\begin{aligned}0 = &\sum_{i=1,2}\frac{dp_{Ai}^*}{dw_A}[(p^* - c)\frac{\partial D_{Aj}}{\partial p_{Ai}} + (p^* - c)\frac{\partial D_{Bj}}{\partial p_{Ai}} - (p_{Bj}^3 - c)\frac{\partial D_{Bj}^3}{\partial p_{Ai}}]\\ &+\sum_{i=1,2}\frac{dp_{Bi}^*}{dw_A}[(p^* - c)\frac{\partial D_{Aj}}{\partial p_{Bi}} + (p^* - c)\frac{\partial D_{Bj}}{\partial p_{Bi}} - (p_{Bj}^3 - c)\frac{\partial D_{Bj}^3}{\partial p_{Bi}}].\end{aligned} \tag{20}$$

A.6 Proof of proposition 1

Here we give conditions for β to be positive.

We use the following notations: $\pi^i(p_{Ai}, p_{Bi}, p_{Aj}, p_{Bj})$ is the profit of retailer i as a function of the four prices, and the derivatives are computed with the variables in this order. Consider $\beta = (p^* - c)(\frac{\partial D_{Ai}}{\partial p_{Bj}} + \frac{\partial D_{Bi}}{\partial p_{Bj}}) - (p_{Bi}^3 - c)\frac{\partial D_{Bi}^3}{\partial p_{Bj}}$.

- First note that $p_{Bi}^3 > p^* \Leftarrow \pi_{12}^j > 0$. Indeed, given the equilibrium prices set by the competitor (p^*_{Aj}, p^*_{Bj}), retailer i sets p^3_{Bi} to solve:

$$\begin{aligned} &\underset{p_{Bi}}{Max} \pi^i(p_{Ai} = +\infty, p_{Bi}, p^*_{Aj}, p^*_{Bj}) \\ \Leftrightarrow\ &\pi^i_2(p_{Ai} = +\infty, p_{Bi}, p^*_{Aj}, p^*_{Bj}) = 0. \end{aligned}$$

Similarly, retailer i sets p^*_{Bi} to solve:

$$\begin{aligned} &\underset{p_{Bi}}{Max} \pi^i(p_{Ai} = p^*_{Ai}, p_{Bi}, p^*_{Aj}, p^*_{Bj}) \\ \Leftrightarrow\ &\pi^i_2(p^*_{Ai}, p_{Bi}, p^*_{Aj}, p^*_{Bj}) = 0. \end{aligned}$$

The implicit theorem functions yields:

$$\frac{dp_{Bi}}{dp_{Ai}} = -\frac{\pi^i_{21}}{\pi^i_{22}}$$

where $\pi^i_{22} < 0$ and $\pi^i_{21} \geq 0$ by Assumption (3), we have $p^3_{Bi} \geq p^*$.

- Consider now β. We have

$$\begin{aligned} \beta &= \pi^i_4(p^*_{Ai}, p^*_{Bi}, p^*_{Aj}, p^*_{Bj}) - \pi^i_4(p_{Ai} = +\infty, p^3_{Bi}, p^*_{Aj}, p^*_{Bj}) \\ &\geq \pi^i_4(p^*_{Ai}, p^*_{Bi}, p^*_{Aj}, p^*_{Bj}) - \pi^i_4(p^*_{Ai}, p^3_{Bi}, p^*_{Aj}, p^*_{Bj}) \text{ if } \pi^i_{14} \leq 0 \\ &\geq \underbrace{\pi^i_4(p^*_{Ai}, p^*_{Bi}, p^*_{Aj}, p^*_{Bj}) - \pi^i_4(p^*_{Ai}, p^*_{Bi}, p^*_{Aj}, p^*_{Bj})}_{0} \text{ if } \pi^i_{24} \geq 0 \end{aligned}$$

Hence a sufficient condition is $\pi^i_{14} \leq 0$ and $\pi^i_{24} \geq 0$. Note that in the linear demand case we have $\pi^i_{14} = \pi^i_{24} = 0$. In that case $\beta = 0$ but $\alpha > 0$.

A.7 Proof of proposition 2

In a linear demand example:

$$q_{Ki} = \frac{1-a-b+ab-p_{Ki}+ap_{Li}+bp_{Kj}-abp_{Lj}}{(1-a^2)(1-b^2)} \tag{21}$$

where $a \in [0,1]$ is a parameter that represents the competition among products and $b \in [0,1]$ is a parameter that represents the competition among retailers.

We obtain:

$$\frac{dp^*_{Ai}}{dw_A} = \frac{1}{2-b} \qquad\qquad \frac{dp^*_{Bi}}{dw_A} = 0$$

B Data

B.1 Construction of the Brand Type Variable

C Test of the Common Trend Hypothesis

One requirement of the DID approach is that the outcomes of the affected and comparison groups follow parallel trends. Although, we cannot assert that, absent the law, prices would have evolved identically in the affected and comparison groups, it is possible to check the "common trend" hypothesis at least for the pre-LME period. Following Ashenfelter, Hosken, and Weinberg (2013), we conduct a formal statistical test that is based on the estimation of the following (log) linear model:

$$\ln\left(P_{ijt}\right) = \alpha + \sum_{l=2}^{59} \varphi_l Month_t^l + \sum_{l=2}^{59} \gamma_l Month_t^l \times T_{ij} + \delta T_{ij} + \mu_{ij} + \varepsilon_{ijt} \tag{22}$$

where T_{ij} is a dummy variable equal to one if the chain-product pair ij belongs to the affected group. The coefficients γ_l measure the price deviation of the affected products from the average price of the comparison products in each month. Table 9 reports the estimated coefficients $\widehat{\gamma}_l$ and the associated t-statistics. We then fit the linear trend of the estimated interaction terms corresponding to the pre-LME time periods, and we test whether the estimated slope is statistically different from zero. We were not able to reject the null hypothesis at a 10% significance level (p-value=0.2130), which indicates that product prices of the affected group do not deviate significantly from those of the comparison group in the pre-LME period.

The conclusion of the test can be reinforced by looking at the significance of each coefficient of the interaction terms in the pre-LME time periods. Overall, we find that the monthly price deviations of the affected group with comparison to the comparison group are rarely significant at the 5% level. However, in the seven months preceding the law one notes that national brand prices deviate significantly from private label prices. The divergence in trends between the two types of product coincides with the sudden rise of agricultural commodities at the end of 2007. One explanation is that the rise of input prices has been transmitted differently in final prices depending on the type of products (i.e., cost-pass through changed). That is, national brands have less passed to consumers the rise of input prices. In order to test whether our results are sensitive to the inclusion of these months, we re-estimate the model by excluding them from the sample. We obtain quantitatively similar results that are still statistically significant.

Table 9: Monthly Price Deviations of Affected Products

Dependent variable: (log) of monthly average price		
Interaction terms	Coef.	t-stat
Treatment × Month2	-0.0023	(-1.00)
Treatment × Month3	0.0015	(0.57)
Treatment × Month4	0.0064	(1.35)
Treatment × Month5	0.0056*	(1.83)
Treatment × Month6	0.0058	(1.38)
Treatment × Month7	0.0105**	(2.10)
Treatment × Month8	0.0129**	(2.34)
Treatment × Month9	0.0068	(1.42)
Treatment × Month10	0.0064	(1.22)
Treatment × Month11	0.0043	(0.81)
Treatment × Month12	0.0008	(0.15)
Treatment × Month13	0.0076	(1.33)
Treatment × Month14	-0.0083**	(-1.96)
Treatment × Month15	-0.0063	(-1.35)
Treatment × Month16	-0.0122***	(-3.07)
Treatment × Month17	-0.0078	(-1.36)
Treatment × Month18	-0.0126***	(-2.72)
Treatment × Month19	-0.0106**	(-2.30)
Treatment × Month20	-0.0053	(-0.86)
Treatment × Month21	-0.0060	(-1.01)
Treatment × Month22	-0.0142***	(-2.95)
Treatment × Month23	-0.0049	(-0.91)
Treatment × Month24	-0.0106**	(-2.05)
Treatment × Month25	-0.0163***	(-3.46)
Treatment × Month26	-0.0056	(-1.12)
Treatment × Month27	-0.0139***	(-3.33)
Treatment × Month28	-0.0147***	(-3.51)
Treatment × Month29	-0.0133***	(-2.67)
Treatment × Month30	-0.0122**	(-2.47)
Treatment × Month31	-0.0141***	(-2.92)
Treatment × Month32	-0.0190***	(-3.69)
Treatment × Month33	-0.0181***	(-2.90)
Treatment × Month40	-0.0448***	(-8.32)
Treatment × Month41	-0.0444***	(-8.22)
Treatment × Month42	-0.0419***	(-7.24)
Treatment × Month43	-0.0428***	(-6.87)
Treatment × Month44	-0.0351***	(-6.09)
Treatment × Month45	-0.0405***	(-6.89)
Treatment × Month46	-0.0378***	(-6.18)
Treatment × Month47	-0.0262***	(-3.85)
Treatment × Month48	-0.0337***	(-5.48)
Treatment × Month49	-0.0365***	(-5.75)
Treatment × Month50	-0.0319***	(-5.26)
Treatment × Month51	-0.0314***	(-5.06)
Treatment × Month52	-0.0259***	(-3.86)
Treatment × Month53	-0.0285***	(-4.84)
Treatment × Month54	-0.0262***	(-4.39)
Treatment × Month55	-0.0200***	(-3.16)
Treatment × Month56	-0.0180***	(-2.71)
Treatment × Month57	-0.0121**	(-2.10)
Treatment × Month58	-0.0105*	(-1.70)
Treatment × Month59	-0.0069	(-1.09)
Treatment × Month60	-0.0019	(-0.29)
Treatment × Month61	-0.0097	(-1.54)
Treatment × Month62	-0.0079	(-1.26)
Treatment × Month63	-0.0107*	(-1.71)
Treatment × Month64	-0.0128**	(-2.14)
Treatment × Month65	-0.0059	(-0.92)
Chain-Product FE	Yes	